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THE ROLE OF ASSET PRICES IN THE FORMULATION OF MONETARY POLICY

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Introduction

Since the early 1980s, the behaviour of asset prices has posed a continuing concern to central banks in their formation of monetary policy as well as a challenge to researchers attempting to explain and interpret the behaviour of asset prices. For instance, the world-wide collapse in equity prices in 1987, the property price cycles in several industrial countries during the second half of the 1980s and the sharp decline in bond prices in 1994 were all to a large extent unexpected and may have established a new asset price environment.

These concerns and challenges were also present last year. While inflation in most industrial countries either declined or remained stable within a 1-3% range, both bond and equity markets recorded substantial gains and exchange rates also moved quite strongly in several countries with floating exchange rates.

This combination of low and stable goods and service price inflation and pronounced movements in asset prices raises the issue of whether, and if so how, monetary policy should react. This is a complex issue, in particular in the current environment where asset price changes in most countries have been confined to financial assets while property prices have been largely stable. To further explore this and related issues, the topic of the Central Bank Economists' Meeting held at the BIS on 28th and 29th October, 1997 was chosen to be:

"The role of asset prices in the formulation of monetary policy".

The 15 papers submitted by the participating central banks were presented and discussed in four sessions, covering the information content of asset prices, the determinants of asset prices, the role of asset prices in the transmission mechanism and asset prices and monetary policy. The rest of this introduction summarises the papers in the order in which they were presented and concludes with a brief overview of the major issues, based on the general discussion during a final session.

Session 1: The information content of asset prices

The paper by *A. Côtè and J.-F. Fillion (Bank of Canada)* reviews the results from an extensive research project on the information content of the term structure of interest rates and its use in the conduct of monetary policy by the Bank of Canada. The paper starts from the observation that there is a tight relationship between the long-short spread and future economic growth in Canada, and uses a small theoretical model to explore why this correlation arises. There are two competing explanations as it may stem either from the effects of monetary policy, or it could reflect endogenous responses of the term structure to expected future changes in economic growth and inflation. Turning to the data, the authors argue that, empirically, the first interpretation seems correct as the peak correlation occurs with a lead of 4-6 quarters, which corresponds to the likely peak effect in the transmission mechanism. Moreover, among the components of demand, the correlation is strongest for durable consumption. The paper also notes that the highest explanatory power of the long-short spread is reached at a longer forecast horizon for inflation than for real output, which is consistent with the chain of causality running from the spread to real economic activity and then to inflation. There are, nevertheless some puzzles. In particular, the term structure has systematically overpredicted output in Canada in the most recent years. Part of this overprediction is probably related to the rise in the risk premium on Canadian bonds in the 1990s but other factors may also be at work.

Next the authors review work on the expectations hypothesis of the term structure of interest rates, concluding that a time-varying term premium is present so that the theory is rejected.

However, so far it has been difficult to find what factors explain movement in the term premium (except the volatility of interest rates). To investigate this, the paper ends with a discussion of the role of the public debt in the determination of long Canadian interest rates. The empirical results lend some credence to the view that increasing public debt in the 1990s has led to "indebtedness premia" in the term structure.

The paper by *D. Domanski and M. Kremer (Deutsche Bundesbank)* first reviews the standard rational valuation formula (RVF) for long-term interest rates and the dividend yield. The authors then empirically assess the question whether the dividend yield predicts future market returns (as implied by the RVF for stocks) and find, similar to results for the United States, that returns are predictable. Moreover, there is some evidence that the dividend yield also predicts future inflation. The information content of the current term spread for future changes in short-term interest rates is also examined and it appears that the medium-term segment of the term structure has most forecasting power for future interest rates, although the expectations hypothesis is rejected for this segment. The predictive content for future inflation is also strongest in the medium-term segment. Finally, the paper discusses the implications for monetary policy. The authors interpret the results as showing that by providing a nominal anchor and implementing policy in a transparent way that reduces short-rate surprises and market volatility, the Bundesbank can facilitate the formation of private sector expectations. However, the usefulness of asset prices as monetary policy indicators seems to be limited to the extent that forecast errors are rather large. Moreover, to avoid circularity problems it is still necessary that the central bank provides an external anchor.

The paper by *T. Jordan (Swiss National Bank)* uses VAR models for variables in levels to assess the predictive performance of long-term interest rates, the slope of the term structure, an effective nominal exchange rate, a stock index and a monetary aggregate (which is M2 rather than the monetary base) with respect to future output and inflation. The results for both in and out-of-sample analyses consistently suggest that money and the exchange rate have the strongest predictive content, with money somewhat better for output and the exchange rate for prices. The paper also tests whether there have been shifts over time in the forecasting ability of the different VAR systems. In this regard, it finds that the predictive ability of money has maintained its superiority even though output has become more difficult to forecast. In contrast, the performance of the exchange rate has eroded over time.

In their paper on the information value of financial asset prices in Spain, *F. Alonso, J. Ayuso and J. Martínez-Pagés (Bank of Spain)* start by testing the predictive performance of a range of asset prices with respect to future inflation, output and three-month interest rates. According to the empirical evidence, these variables do not seem to add any information beyond what is contained in past values of the variables themselves, though it does appear that indicators based on the term structure work slightly better. Given this result, the authors use probit models to see whether financial variables are useful for predicting qualitative variables, such as accelerations of inflation, slowdowns in output growth or a tighter monetary policy. In this case, financial variables do add information to that contained in past values of the dependent variable. Finally, the authors review recent research at the Bank of Spain on whether financial variables are useful measures of expectations held by investors. The research suggests that movements in long interest rates are largely due to shifting inflation expectations (i.e. real interest rate and inflation risk premia are relatively stable) and that short-term interest rates are good measures of expected future short rates (i.e. term premia are small).

Session 2: Determinants of asset prices

The objective of the paper by *M. Dombrecht and R. Wouters (National Bank of Belgium)* is to explore the information content of the term structure of interest rates. Because the intermediate target of monetary policy in Belgium is a stable exchange rate against the Deutsche mark and the two countries are close linked through trade, the information content of the Belgian term structure cannot be isolated without extending the model to include the German term structure. This close link is already evident in the first part of the paper which estimates the response of market rates to changes in official rates. For both countries there is a relatively high degree of persistence, suggesting that markets see official rates as mean reverting; however, persistence is higher in Germany than in Belgium, even when the period around the 1993 exchange rate turmoil is left out. Moreover, the estimated reaction of long rates to changes in official rates imply that changes in official rates are mostly unanticipated in Belgium; in contrast, such changes are to a large extent anticipated in Germany, suggesting that the credibility of monetary policy is very high.

The authors next turn to correlations between term spreads and future inflation and real growth, starting with a discussion of how the leading character of the term spread should be interpreted and how reliable the spread is as a source of information for monetary policy. Having outlined several channels of interaction and underlined the importance of identifying endogenous reactions of monetary policy, or reactions as perceived by market participants, to expected changes in inflation and output growth, the authors present several empirical tests. According to causality tests, the short-term rate has a much stronger predictive impact than the long rate, suggesting that most of the correlations can be attributed to policy actions. Complementary evidence obtained from estimating structural VAR systems corroborates this finding as most of the covariance between the term structure and future inflation can be related to short-term interest rate shocks whereas inflation shocks provide only a small contribution. All in all, the authors conclude that prudence is required when using the yield curve as a forward-looking indicator for monetary policy formulation.

The paper by *G. Grande, A. Locarno and M. Massa (Bank of Italy)* addresses the issue as to why simple regressions of stock market returns suggest that stock markets provide a poor hedge against inflation. They do so by constructing and estimating a model in which the way in which inflation affects stock prices depends on the monetary policy regime. Following a brief review of the literature, the authors first estimate the Fama and Schwert (1977) equation for various assets in Italy and find that interest rates at various maturities provide an incomplete hedge against expected inflation and no hedge at all against unexpected inflation. Moreover, stock returns are insensitive to expected as well as unexpected changes in inflation.

Since the empirical analysis also indicates that parameters have not been stable over time, the authors next use a methodological framework proposed by Campbell which decomposes unexpected excess returns into changing expectations concerning future dividends, short-term real interest rates, inflation and required excess returns. Moreover, market expectations of these variables are allowed to vary depending on some unobserved monetary policy regime. For example, when the central bank is perceived to conduct a tight monetary policy, even a small positive increase in the expectations of inflation may induce markets to expect a strong monetary policy reaction. Conversely, when policy is not strongly committed against inflation, a surge in inflation does not necessarily mean lower future economic growth and the Fisher hypothesis may hold. However, unlike the previous literature, the paper does not exogenously impose policy shifts but derive the prevailing regimes from a model that allow for endogenous and data dependent changes.

According to the empirical results, the data clearly indicate a significant regime shift around 1987-88, attributable to a change in the operation of monetary policy, combined with institutional changes and adjustments of the wage indexation scheme. The empirical tests reject the hypothesis of a one-to-one relationship between short-term interest rates and expected inflation in Italy, though short rates partly incorporate inflation forecasts. The results further show that equities have not significantly outperformed government securities as hedges against expected inflation over the last two decades.

In a paper on Japanese share prices, *S. Uemura and T. Kimura (Bank of Japan)* demonstrate that equity returns have been good predictors (in the Granger sense) of CPI inflation and various measures of real activity in the 1970-97 period. The paper further documents that stock and land prices have moved very closely together. Next, an analysis of the sources of Japanese equity price movements is presented, using a simple model which expresses the risk premium as a function of the P/E ratio, long-term interest rates and the expected growth of earnings (or dividends). By adjusting the P/E ratio for cross holdings and cyclical factors and using corporate growth forecasts as a measure for earnings growth, the authors are able to solve for the risk premia and interpret movements in Japanese stock prices. One interesting finding in this context is that the current low level of stock prices reflects high-risk premia and low expectations of earnings growth.

The authors then attempt to explain variations in the calculated risk premium. They do so by regressing it on past inflation, the standard deviation of past industrial production growth, the bankruptcy rate of firms and the spread between CDs and T-bills, finding that these variables explain a large fraction of the computed risk premium. Finally, in reviewing some structural changes in the stock market, the authors document that institutional and foreign investors have gradually increased their role in the market.

In the first part of his paper on Swedish stock prices, *P. Sellin (Bank of Sweden)* calculates fundamental values on the assumption that dividends obey a random walk, so that the fundamental stock price is a fixed multiple of dividends. From this model, it appears that stock prices were undervalued between 1919-50, appropriately valued between 1950-81, and overvalued 1981-96. The paper then uses a theoretical general equilibrium asset price model to discuss the impact of monetary policy on equity prices. According to this second model, a tighter monetary policy should increase equity prices by reducing inflation and increasing the real value of future dividends. However, the author also argues that other asset price models may lead to the opposite result, concluding that, ultimately, the effects of monetary policy on stock prices in Sweden has to be determined empirically. Applying a GARCH model and using various dummy variables as proxies for the potential impact of speeches by the Governor and the Deputy Governor, the inflation reports as well as changes in policy interest rates, there is some evidence that a tightening of monetary policy leads to lower stock prices and higher long bond yields.

The paper by *S. Hayes, C. Salmon and S. Yadav (Bank of England)* first attempts to explain the recent increase in UK equity prices by developments in the equity risk premium and expectations of future dividend growth. To determine the risk premium the authors rely on the CAPM model, implemented by an EGARCH-M specification which allows for an asymmetric response of the conditional variance to positive and negative shocks. The results for the G-7 countries indicate strong, though variable, correlations of risk premia but do not suggest that the world-wide rise in equity prices can be attributed to lower risk premia. This finding, however, may reflect the assumption that the risk aversion coefficient in the model has been constant over time. The paper next applies the Gordon Growth Model to estimate expectation of future dividend growth. According to the evidence, investors expect future dividends to grow at about twice the rate observed in the recent past. However, this result is also subject to caveats as the model assumes that the equity market is in steady state, which may not be the case. Moreover, it is highly sensitive to estimates of required returns and to the risk that the equity cost of capital may be overestimated.

The authors then turn to the role of asset prices in the transmission mechanism, focusing on balance sheet effects for the corporate sector. The "credit view" of the transmission mechanism suggests that variations in the net worth of corporates will change the premium that some firms pay for external finance, thus adding an additional channel to the monetary transmission mechanism. Indirect evidence indicates that this factor may be of importance in the United Kingdom. In particular, small manufacturing firms experience bigger variations in their external finance premium than large firms and those manufacturing sectors that can be characterised as "small firm sectors" tend to be more sensitive to monetary shocks than "large firm sectors". The last two sections of the paper report on preliminary and ongoing work at the Bank on the information content of equity prices with respect to future output and inflation and on the use of option prices to gain ex ante information about future market changes. Based on the preliminary results obtained from VAR models, it appears that equities contain information about the real economy whereas only a small part of the variation in inflation can be attributed to real equity returns. However, more precise results may be obtained through disaggregation and sub-period analyses as well as by complementing such work by extracting expectations of future dividends and discount rates from equity prices. Regarding option prices, the authors note that more work is required to determine whether moments of the implied distributions can predict future market changes and the expenditure behaviour of consumers and firms.

Session 3: The role of asset prices in the transmission mechanism

The first part of the paper by *H. Glück and R. Mader (Austrian National Bank)* analyses the ability of term spreads to predict future industrial production and inflation. It appears, however, that term spreads have little predictive power for Austria and the inclusion of stock prices does not change this result. The authors list several reasons why spreads and equity prices contain little information about future economic conditions. First, the sample period is short and Austrian interest rates have hardly responded to market forces before deregulation in the late 1980s. Moreover, institutional investors, who arguably react more rapidly to shifting expectations of the future, play a relatively limited role in Austrian financial markets. Most importantly, monetary policy in Austria has been based on a fixed exchange rate against the Deutsche mark, implying that interest rates have been determined by economic conditions in Germany which are imperfectly correlated with those in Austria. The second part of the paper briefly summarises and reviews earlier work on the role of wealth in the transmission mechanism in Austria. It highlights various channels through which wealth effects may affect economic conditions, including consumption, investment and inflation expectations as well as bank lending and money demand. Even allowing for recent developments, none of the channels have, so far, been very important in Austria and the paper concludes by discussing various reasons for this.

The paper by *P. Jaillet and P. Sicsic (Bank of France)* first focuses on the influence of equity and house prices on consumption. The authors note that French households hold directly about 30% of outstanding shares, compared with 64% for US households, whereas the distribution across individuals is more concentrated in France than in the United States. The paper next presents some econometric tests that fail to find any link between equity prices and the growth of consumption. An increase in bank lending and a boom in commercial property prices are documented in the following section whereas no housing price increase is observed for France as a whole. The authors then turn to the issue of whether asset prices condition the responses of bank interest rates to monetary policy measures, using panel data on bank lending. According to the empirical tests presented, there is no evidence that changes in banks' balance sheets affect the extent to which they adjust lending rates to changes in market rates; in fact, the degree of inertia may have fallen over time. In the final part of the paper, the authors turn to the implications of asset price movements for monetary policy. They first note that property prices, in particular, are subject to measurement errors and thus not a reliable source of information. Moreover, monetary policy is not the right instrument for dealing with a potential asset price bubble and targeting asset prices entails a risk of generating pro-cyclical interest rate changes and, in particular, might pose a danger to the ultimate goal of price stability.

In their paper on "Asset market hangovers and economic growth", *M. Higgins and C. Osler (Federal Reserve Bank of New York)* first develop an empirical model to explain house price developments in each of the fifty US states from 1973 to 1996 and to detect whether bubbles or misalignments were present. Based on a standard present discounted value model, their regression includes state disposable income, employment, construction costs and real mortgage interest rates as the fundamental factors, while non-fundamental factors comprise measures of credit availability, overbidding and the lagged ratio of real house prices to real disposable income, interpreted as an affordability index. While most of the fundamental factors have the expected sign and are significant, only the affordability index is significant among the non-fundamental factors. The authors interpret this as evidence that credit availability (or collateral effects) is not very important in determining house prices. Using this model the authors then derive a measure of misalignment, defined as the total departure of house prices from the fundamental factors They find that four of the nine US census regions exhibit a non-fundamental component, consistent with the notion of a bubble in that both the rise and the subsequent fall of house prices are fairly monotonic.

The authors then examine whether the unexplained house price booms and busts have affected housing investment. Starting from the neoclassical theory of investment, they derive an estimable equation which explains housing authorisations (their measure of housing investment) in terms of lagged house prices, per capital income, the unemployment rate, the user cost of capital and mortgage originations. The estimated parameters point to a reasonably strong link between house prices and investment, which the authors interpret as a Tobin's q effect. As a result, housing investment appears to have been significantly affected in regions that have experienced bubbles. The final section briefly discusses policy implications. Although it is acknowledged that monetary policy could be used to mitigate the effects of housing price bubbles, the authors have reservations. The bubbles found are regionally concentrated and it is hard to define bubbles ex ante. Moreover, if monetary policy is used, it might become less transparent. An alternative would be tax policies or regulations, but these policies also face problems if bubbles are regional and hard to define ex ante. Moreover, an interesting corollary of the findings is that monetary policy and/or differences in the regulatory environment cannot explain why bubbles arose in particular states.

Session 4: Asset prices and monetary policy

The paper by *C. Kent and P. Lowe (Reserve Bank of Australia)* starts by reviewing two conclusions from the existing literature on asset prices: (i) to assess the effects of asset prices, it is necessary to understand the sources of asset price movements; (ii) prices of assets that are used as collateral (e.g., real property) are of greater importance for policy than those that are not (e.g., equities). It next addresses the question of how monetary policy should respond to asset price bubbles, using a simple model. The most important conclusion from the model is that there may be circumstances where monetary policy should be tightened in response to an emerging asset-price bubble, in order to burst the bubble before it becomes too large, even though this means that expected inflation is below target in the short run. Such a policy is optimal because it can help to avoid extreme longer-term effects of a larger asset-price bubble and its eventual collapse. The authors then review the implications of low inflation for asset prices. They suggest that in a low-inflation environment asset-price bubbles are less likely, although, if they occur, they may be more costly. If inflation is low, a fall in the real asset prices will largely come through nominal price declines, which are likely to be particularly adverse to economic activity.

The authors then go on to review the movements in credit and asset prices in Australia in the period 1966-97, concluding that recessions coincide with falls in property prices. Moreover, credit growth is closely tied to property prices, but appears unrelated to equity prices. The paper presents econometric estimates of the relationship between asset price swings and real growth, concluding that fluctuations in real property prices are strongly related to output growth, notably in periods when property prices are falling. However, these regressions are not necessarily structural, and the significance of asset prices may reflect that they are forward-looking variables. It is also found that asset prices do not seem to influence inflation, except through the output gap.

The lessons drawn for monetary policy are, first, that credit and property-price cycles go hand in hand. Second, monetary policy can burst a bubble in property prices and, in some circumstances, it makes sense to do so, whereas it makes less sense to burst equity bubbles. Third, while financial liberalisation is important, credit and property cycles occurred in Australia before deregulation; regulation mainly determines who extends credit, but not necessarily how much credit is extended. Fourth, low inflation can make bubbles more costly, since the adjustment to a bursting bubble requires nominal asset prices to fall.

The paper by *J. Capel and A. Houben (Netherlands Bank)* defines asset price inflation as occurring when asset prices are rising and exceed fundamentals. Using some simple back-of-the-envelope calculations based on various models in the literature, the authors conclude that Dutch equity prices are likely to be too high whereas current property prices do not seem excessive, once various fundamentals, such demographic changes and quality improvements, are taken into account. The paper goes on to argue that the major risk with asset inflation stems from the consequences of the "bubble bursting". The paper reviews the implications for the banking sector in the Netherlands and judges the effects of a fall in equity prices to be small. Similarly the impact on consumption and investment is found to be rather moderate, as households hold a relatively small proportion of equities and investment is mostly financed by bank credit or own funds. In contrast, the effects of a fall in property prices would be more worrisome as housing wealth is relatively high and evenly distributed.

In the last section of the paper, the authors turn to the question of the monetary policy implications of asset price inflation. It is difficult to identify the sources of asset price changes. Some may be real or fundamental changes, such as changes in tax codes or variations in risk premia while others are non-fundamental. Consequently, it is also hard to say what the central bank should "lean against". Furthermore, the idea of stabilising an index of goods and asset prices is impractical. Nevertheless, asset prices may condition monetary policy through their direct and indirect effects on CPI inflation. Thus rising asset prices are likely to increase CPI inflation by increasing costs of capital goods and housing and by affecting expenditure. Confidence effects could also play a role, as could self-reinforcing effects through bank credit expansion. In the latter case, however, a tightening of supervisory policy seems more appropriate, as monetary policy is a very blunt instrument. Moreover, in an economy operating under fixed exchange rates there is little scope to gear policy to asset prices. Overall, the empirical evidence for the Netherlands suggests that the link between asset prices and inflation is not very tight.

The paper by *V. Reinhart (Federal Reserve Board)* basically consists of three parts. In the first part, an attempt is made to determine whether equity prices are related to macroeconomic fundamentals and to assess whether current equity prices are overvalued. To this end, the author estimates an error-correction model based on the Gordon growth equation, which explains the price-earnings ratio in terms of inflation expectations, the rate of unemployment and six alternative interest rates. According to the estimates, the earnings-price ratio moves positively, and mostly one for one, with long-term interest rates and negatively with the rate of unemployment, whereas expected inflation does not seem to have any influence. Next, the estimated long-run relationship is used to derive a measure of deviations from fundamentals. This suggests that current stock prices may be overvalued, although far less than before 1987. More generally, the distribution of deviations shows spikes at correct valuations and at significant undervaluations.

The second part of the paper discusses the effect of equity values and the term spread on spending. Using a simple error-correction model, estimated simultaneously for GDP, private consumption, business fixed investment and imports as the dependent variable and separately for 1973-84 and 1985-96. From the estimation results, financial variables do not appear to have exerted a systematic influence on spending prior to 1985. In contrast, significant elasticities are identified for the post-1985 period, ranging from ⅓ for real GDP and real consumption over ⅔ for real investment to unity for real imports. Moreover, business investment seems to respond more strongly to changes in the term spread than the other demand components.

Against this background, the third part of the paper discusses whether monetary policy should react to equity prices. Using a simple theoretical model of the economy to describe the interaction between equity prices and interest rates, the author demonstrates that assigning a more important role to equity prices in the setting of monetary policy than what they imply for forecasting spending and inflation is discouraging. First, a greater policy responsiveness to the *level* of equity prices could increase the effect of news on evaluation and thus might lead to more volatility. Second, a more active policy response to *changes* in equity prices sets up a feedback loop which raises the reaction of equity prices to policy misalignments and could be destabilising.

Summary of discussion[1]

The role of asset prices, whether quotes on government securities, equities or foreign exchange rates, in the conduct of monetary policy probably depends on circumstances. In normal times, policy makers might consult market quotes as inputs to structural forecasts, as indicators of

1 The following is essentially based on the summary by *Mr. Reinhart (Federal Reserve Board).*

future activity and inflation, or as explicit targets. At times of stress, however, the behaviour of asset prices might give reasons for treating the near-term outlook as more uncertain or as a call for special action.

Several participants noted various entry points of financial market prices along the traditional monetary policy transmission mechanism, in particular how changes in the policy rate get reflected along the term structure and how changes in government yields get translated into private rates, equity prices and exchange rates. As noted in the paper by *Mr. Reinhart*, current values of market yields are, with the exception of equity prices, fairly close to fundamentals. This suggests that if there is an asset bubble, it is not widely inflating. If there is a problem with equity prices, it is a problem of *relative prices* and, with only one national and blunt instrument at its disposal, monetary policy is generally not seen as appropriate for dealing with relative prices.

Further along the transmission mechanism, there is the issue of how changes in financial market prices affect spending. For the United States, long-run reduced-form estimates of these sensitivities show that GDP and its components (consumption, business fixed investment and imports) all depend on the slope of the term structure and equity prices. Moreover, the sensitivities post-1984 are sizeable, though work is required to provide firmer foundations for explaining behaviour as well as a more complete description of dynamics.

The last chain in the transmission mechanism is inflation, which may also be directly influenced by asset prices, notably the exchange rate (for instance, during the last few years, the US inflation has benefited from the appreciation of the dollar) and, perhaps more recently, equity prices. An interesting question in this respect is whether buoyant equity prices, which give managers the ability to compensate workers in the form of options and direct share grants without raising base salaries, may also be imparting some restraint.

As regards indicators serving as an independent check on judgmental or model-based forecasts, various financial market prices have been offered, for instance, the slope of the yield curve and the level or growth of equity prices. Emphasising such indicators is not necessarily a rejection of a traditional structural approach to understanding the economy. Rather simple rules of thumb may be useful as a cross-check on that understanding and, in a world where it is important for central banks to explain their actions, as pedagogic devices.

Some participants expressed concern that the messages obtained from markets were asymmetric in that they signalled the onset of recession more reliably than expansion. One reason may be that it is easier to recognise signals in the former than in the latter case. A yield curve that is inverted is an obvious signal that economic conditions are at variance with sustained expansion. A similar process was at work in the United States in early 1994, when the Chairman of the Federal Reserve Board pointed to a real federal funds rate of zero as a reason to tighten policy. In such an extreme circumstance, there is a lot of information that can easily be conveyed. In general, however, knowledge of the determinants and behaviour of the term structure is sufficiently primitive that little is known about when the yield curve is "too steep". Similarly, using the real federal funds rate to motivate and to explain policy actions gets more difficult when the issue concerns determining and settling at an "equilibrium" federal funds rate.

Regardless of whether asset prices are used as part of a structural interpretation of the economy or as a reduced-form indicator, there are reasons for central banks to monitor the behaviour of financial markets. However, asset prices are a means to an end, not the end itself. It is a completely different matter to use an asset price as a target for policy, though it is not an uncommon suggestion. In many countries, monetary policy uses the nominal exchange rate as the normal anchor. Perhaps more relevant to the topic of this meeting, policy makers have, at various times, also indicated that monetary policy be guided by commodity prices or the slope of the yield curve. Among the arguments for using market quotes in the policy process have been that such prices are forward looking (often over long horizons), timely, cheap to collect and rarely revised.

However, the problems with such an approach are considerable. First, as mentioned by several participants, there is no reason to expect that historical relationships between an asset price and the ultimate target variable will remain stable if a central banks tries to exploit them. As a general rule, because asset prices depend importantly on expectations, they are not likely to be stable functions of any small set of variable through time. Second, if a systematic policy response does reduce the volatility of one asset price, it probably just means that pressures have been shifted to other prices. This follows the logic of Dornbusch's overshooting model: when goods prices are sticky, the flexible price (i.e. the exchange rate) has to move by more than when goods prices are flexible. Third, the systematic response of monetary policy to an asset price need not make the asset price more stable. For example, as mentioned in the papers by *Messrs. Domanski and Kremer and Mr. Reinhart*, if a central bank moves the short rate in response to past changes in equity prices, current and future prices tend to become more volatile. That is, if a central bank were to act like a feedback trader, it adds to volatility. However, the notion of "leaning against the wind" of price movements is that sort of feedback rule. Finally, interesting theoretical possibilities follow when a central bank sets its short rate to maintain some desired slope to the term structure. For instance, if inflation is backward looking, then this rule would make inflation a random walk. Essentially, the central bank accommodates the current rate of inflation and if a shock pushes inflation up or down, the new rate gets built into longer-term nominal rates and policies must realign the short rates accordingly. Conversely, if inflation is forward looking, then this rule makes inflation indeterminant. Any initial judgement by market participants about the appropriate level of inflation will call forth a policy response so there is nothing to anchor the system.

Monetary policy makers might be concerned about equity or property prices straying above fundamentals due to potential adverse effects as asset prices rise or as they fall back to earth. When prices move above fundamentals, relative prices are misaligned, dictating some misallocation of resources. Households might be consuming out of their paper wealth and firms buying capital based on inflated market-to-book values. Moreover, favourable leverage ratios and receptive capital markets may induce households and firms to take on debt and new firms to start up. Nonetheless, to the extent that a specific asset price inflation does not signal a more general misalignment, suggesting that policy was creating excess liquidity, there is little that the central bank can do. If the relative price has sufficient macroeconomic consequences so as to warrant policy action, that action would be based on achieving a desired macro-outcome and not on managing a relative price.

Policy makers might also be concerned about increases in stock prices, fearing that there might be time dependence to overvaluation, on the grounds that the longer prices stray above fundamentals, the further they stray and the harder they will fall. One example is a rational bubble, which pushes prices further above fundamentals the longer it lasts, as the continuing rise in equity prices above fundamentals provide investors sufficient excess returns to compensate them (in an expected value sense) for the decline when the bubble bursts. Presumably, as explained in the paper by *Messrs. Kent and Lowe*, if the bubble bursts sooner, the less adverse will be the misallocation of resources during the transition and the less severe the systemic strain when prices drop

The stresses induced by sudden and acute changes in asset prices may influence monetary policy making in two ways. For one, a larger realignment in prices may be a reason to treat the near-term outlook as more uncertain and for monetary policy makers to revisit the foundations of their own forecasts. For instance, most of the models derived from the data are linear and large price changes may test the robustness of such relationships. In addition, large swings in asset prices may not be due to market dynamics alone. There may be factors related to fundamentals, such as market participants changing their outlook in a way that triggers sudden and large changes in key financial market prices.

A large realignment of prices may also require special policy action. For instance, equity prices might be a source of concern on their way down because of systemic risks, knock-on effects on spending and confidence and the risk of subsequent undershooting. With regard to systemic risk, mechanisms are well developed for dealing with such problems, including the discount window and the willingness to add ample reserves at times of stress.

Participants in the meeting

Australia:	Mr. Philip LOWE Mr. Christopher KENT
Austria:	Mr. Heinz GLÜCK Mr. Richard MADER
Belgium:	Mr. Michel DOMBRECHT Mr. Raf WOUTERS
Canada:	Ms. Agathe CÔTÉ Mr. Jean-François FILLION
France:	Mr. Pierre JAILLET Mr. Pierre SICSIC
Germany:	Mr. Dietrich DOMANSKI Mr. Manfred KREMER
Italy:	Mr. Giuseppe GRANDE Mr. Alberto LOCARNO
Japan:	Mr. Shuichi UEMURA Mr. Takeshi KIMURA
Netherlands:	Mr. Aerdt HOUBEN Mrs. Jeannette CAPEL
Spain:	Mr. Juan AYUSO Mr. Fernando RESTOY
Sweden:	Mr. Ossian EKDAHL Mr. Peter SELLIN
Switzerland:	Mr. Andreas FISCHER Mr. Thomas JORDAN
United Kingdom:	Mr. Chris SALMON Mr. Sanjay YADAV
United States:	Mr. Matthew HIGGINS *(New York)* Mr. Vincent REINHART *(Washington)* Mr. Michael LEAHY *(Washington)*
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The term structure of interest rates and the conduct of monetary policy in Canada

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Introduction

The aim of this paper is to put into perspective the empirical results obtained at the Bank of Canada and elsewhere with regard to the information content of the term structure of interest rates, and to describe how this information is currently used in the conduct of monetary policy in Canada.

There are three main reasons for central banks' interest in financial asset prices. First, since monetary policy influences the economy through the financial markets, the central banks wish to understand the role played by the prices of different financial assets in the monetary policy transmission mechanism. Secondly, financial asset prices may contain useful information for the conduct of monetary policy, irrespective of whether they play an important role in the transmission mechanism. This is because such data may contain more up-to-date information on the economic situation than that otherwise available to the central bank. Moreover, they reflect the expectations of market participants with respect to future economic developments. Since expectations are derived from market transactions, they are often considered to be more representative than the figures obtained from surveys. Finally, changes in financial asset prices can signal market imbalances, which may spill over into the real economy and thus have consequences for monetary policy.

Unfortunately, certain financial data have only been available in Canada for a few years and the markets on which these securities are traded are extremely narrow. This is the case for inflation-indexed bonds and short-term interest rate options. On the other hand, the data making up the term structure of interest rates are readily available and, generally, of good quality. It is partly for this reason that this study is confined to the information content of the term structure. Furthermore, it is a source of information which has been the subject of numerous studies using Canadian data and which has proved to be one of the most conclusive. The existence of a close correlation between the yield curve spread and economic activity raises questions concerning the monetary policy transmission mechanism. This will be the topic of the next section. In Section 2, we examine the way in which the information in the term structure is used at the Bank as part of simple indicator models for real output, inflation and market expectations with regard to future interest rates. Given that the expectations hypothesis of the term structure (EH) plays a dominant role in the analysis of the monetary policy transmission mechanism, Section 3 tests this hypothesis. We then examine the possibility that the presence of a risk premium on internationally-traded Canadian securities, as a result of Canada's high level of indebtedness, partly explains the variability of the term premium and the fairly frequent statistical rejection of the EH. Section 4 describes our main avenues of research.

1. The role of the term structure of interest rates in the monetary policy transmission mechanism

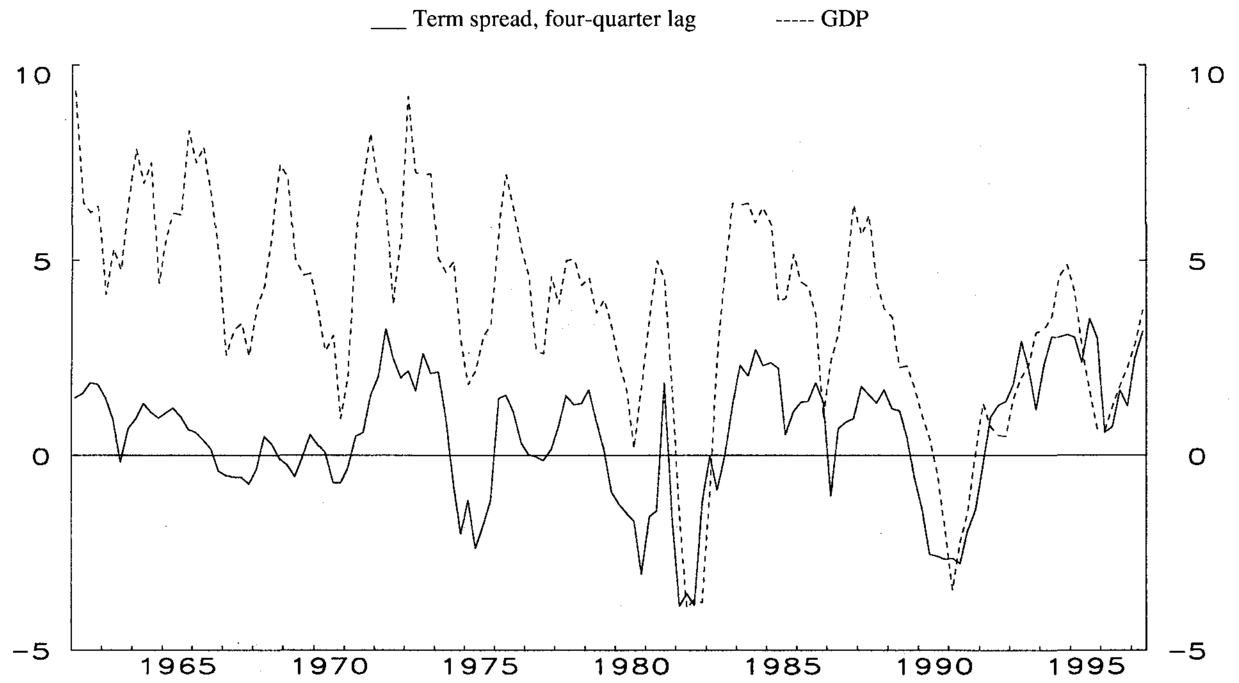
In this section, we review the various hypotheses proposed in the literature to explain the correlation observed between the yield curve spread (for short, the yield spread) and economic activity. We also discuss the empirical results obtained at the Bank, which led the researchers to choose the yield spread as the key monetary variable in the Quarterly Projection Model (QPM).

1.1 Background

Over the last few years, research carried out both at the Bank of Canada and elsewhere has revealed the existence of a strong positive correlation between the yield spread and the subsequent growth of economic activity (Graph 1).[1] In itself, this correlation is not very surprising: according to the traditional model of monetary policy transmission, the central bank affects economic activity through real interest rates. The central banks influences monetary conditions by modifying the amount of liquidity in circulation in the banking system, which has an immediate impact on nominal and real very short-term interest rates. The movements in these rates in turn affect the whole interest rate spectrum and the exchange rate, depending on lenders' and borrowers' expectations with regard to subsequent changes in rates. Since the expectation formation process is a complex one, the response of real long rates cannot be determined a priori, but one would expect it to be smaller than that of short rates since monetary policy measures can have only a temporary effect on real interest rates. In the long run, real interest rates are chiefly determined by expectations as to the productivity of capital and the underlying forces affecting saving globally. In fact, for a small open economy like Canada, world real interest rates provide a stable anchor for domestic real rates.

Graph 1

Lagged yield spread and four-quarter growth rate of real GDP



Thus, the correlation observed between the yield spread and real economic activity could quite simply reflect the endogenous response of these two variables to monetary policy actions. For example, a monetary tightening leads to a narrowing of the yield spread, followed a few quarters later by an economic slowdown. Conversely, a monetary easing results in a widening of the yield spread and faster growth.

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1 See, in particular, Laurent (1988), Cozier and Tkacz (1994), Estrella and Hardouvelis (1991), Hu (1993), Plosser and Rouwenhorst (1994), Estrella and Mishkin (1995, 1997), Haubrich and Dombrosky (1996), Harvey (1997) and Dueker (1997).

What appears less compatible with the traditional point of view is the size of the correlation observed between these two variables and the fact that the yield spread appears to be better than other monetary policy indicators - especially the real interest rate - at forecasting the growth of real economic activity. As Table 1 shows, certain yield spreads are capable, on their own, of explaining approximately 65% of the variance of the future rate of growth (four quarters) of Canadian real GDP in the period 1972-90. The explanatory power of both the level of and changes in short-term interest rates, whether nominal or real, is substantially lower over the same period.

Table 1

Comparison of different interest rate variables in an indicator model of real GDP for the period 1972Q1–1990Q4

$$G4Y_t = a + b(R)_{t-4}$$

R	*b*	$\overline{R}^2$
(1-3 years) – 90 days	1.6 (4.3)	0.38
(3-5 years) – 90 days	1.5 (5.5)	0.50
(5-10 years) – 90 days	1.4 (6.8)	0.58
(10 years+) – 90 days	1.3 (7.9)	0.64
(3-5 years) – (1-3 years)	6.2 (6.7)	0.58
(5-10 years) – (1-3 years)	3.6 (8.5)	0.66
(10 years) – (1-3 years)	2.7 (8.0)	0.66
(5-10 years) - (3-5 years)	7.1 (8.4)	0.62
(+10 years) - (3-5 years)	3.9 (6.7)	0.60
(+10 years) - (5-10 years)	6.0 (3.5)	0.40
4Q moving average (1-day rate)	-0.5 (3.2)	0.31
4Q moving average (real 90-day rate)	-0.3 (2.4)	0.16
Δ_4 4Q moving average (1-day rate)	-0.8 (4.6)	0.38
Δ_4 4Q moving average (real 90-day rate)	-0.7 (3.1)	0.26

Notes: *t*-statistics in parenthesis. $G4Y_t$ = rate of growth of real GDP over four quarters in period *t*. Δ_4 = four-quarter difference operator. 90-day rate refers to commercial paper; other rates correspond to Canadian government bond rates; real 90-day rate is calculated as the nominal rate less the change in the GDP deflator.

According to the traditional point of view, the more closely the path of long rates tracks that of short rates, the greater will be the impact of the initial change of the latter on economic activity. The fact that the yield spread is the best advance indicator of output could be interpreted as signifying the opposite, a result which is obviously not compatible with economic theory and which has prompted researchers to suggest other explanations.

1.2 Suggested explanations

In order to examine the hypotheses put forward in the literature to explain the correlation observed between the yield spread and economic activity, we use the analytical framework of Cozier and Tkacz (1994). It is based on three key hypotheses, which are represented by the following equations:

$$i_i = r_t + E_t \pi_{t+1} \tag{1}$$

$$i_t^k = \frac{1}{k} \sum_{i=0}^{k-1} E_t i_{t+i} + \rho_t^k \tag{2}$$

$$r_t = r^*_t - l_t \tag{3}$$

where i = nominal short interest rate (one period)

r = real interest rate

E_t = expected value on the basis of the information available at time t

π = inflation rate

i_t^k = nominal long interest rate (k periods)

ρ_t^k = term premium

r^* = equilibrium real interest rate

l = liquidity effect of monetary policy

Equation (1) corresponds to the Fisher relationship, which assumes that the nominal interest rate is equal to the sum of the real interest rate and the expected inflation rate. Equation (2) represents the expectations hypothesis of the term structure, according to which the yield on a long-term bond is a weighted average of the expected short rates plus a term premium. Finally, equation (3) states that the real interest rate is made up of two components: the equilibrium real rate, which reflects market forces, and the "disequilibrium" rate, which reflects monetary policy shocks, commonly referred to as the liquidity effect. From these three equations, we can derive the following expression for the yield spread:

$$i_t^k - i_t = \left[\gamma l_t - \frac{1}{k} E_t \sum_{i=1}^{k-1} l_{t+i}\right] - \left[\gamma r^*_t - \frac{1}{k} E_t \sum_{i=1}^{k-1} r^*_{t+i}\right] - \left[\gamma E_t \pi_{t+1} - \frac{1}{k} E_t \sum_{i=1}^{k-1} \pi_{t+1+i}\right] + \rho_t^k \tag{4}$$

where $\gamma = (k-1)/k$.

Equation (4) shows that the yield spread comprises four elements: the liquidity effect of monetary policy via expectations; expectations about changes in the equilibrium real interest rate; expectations concerning future inflation; and the term premium. The fact that the yield spread is negative at cyclical peaks and high and positive during the troughs may therefore reflect one or other of the following factors: countercyclical monetary policy; the cyclical development of the demand for credit, of the return on capital and thus of the equilibrium real interest rate; and the cyclical pattern of inflation, which tends to fall during recessions and rise during periods of expansion.

In order for the yield spread to reflect mainly the monetary policy stance and *bring about* fluctuations in economic activity, its variability must be dominated by liquidity effects. This will be the case if the changes in the equilibrium real rate are either weak or show a high degree of persistence, if inflation expectations show a high degree of persistence and if the term premium is relatively stable over time or is influenced by the liquidity effect. As suggested by Laurent (1988), the yield spread may be a better indicator of monetary policy than the level of interest rates if it allows the monetary component of interest rate changes to be isolated. For this to happen, monetary policy would have to exert a fairly large influence on real short-term interest rates but a relatively small one on real long-term interest rates, which, for their part, would more accurately reflect market equilibrium forces.

On the other hand, certain economists, notably Hu (1993) and Harvey (1997), maintain that the correlation between the yield spread and future output reflects the endogenous response of the term structure of real interest rates to the forecast evolution of economic activity, as predicted by the Capital Asset Pricing Model (CAPM). This model is based on the hypothesis that individuals are risk-averse and try to smooth their consumption over time. If they expect a recession, they will lengthen the term of their investments in order to secure a certain revenue during this period. This substitution of short-term securities with longer-term ones will bring down the price of the former and raise the price of the latter. Thus, the yield curve flattens or is inverted before the economy slows down. Harvey maintains that the same type of reasoning can be applied to the behaviour of firms. If they foresee an economic downturn, they will cut back their long-term investment projects and, as they try to balance the maturity structure of project loans, the supply of long-term securities will fall, which will also lead to an increase in their price and a flattening of the yield curve.

1.3 Empirical results

The three factors mentioned above can all contribute to the predictive power of the yield spread and, in practice, it is not easy to determine which of them is dominant, since expectation variables and equilibrium real rates are not directly observable and all the interest rates are strongly correlated amongst themselves. Estrella and Hardouvelis (1991) and Plosser and Rouwenhorst (1994) conclude that while monetary policy does have a role to play in the predictive power of the yield spread, other factors are important as well. This is because the yield spread remains significant in indicator models which that additional variables to represent monetary policy, such as the level of interest rates or the growth of a monetary aggregate.

At the Bank, we have a certain number of empirical results, which suggest that the predictive power of the yield spread chiefly reflects its role as an indicator of monetary policy:[2]

- the predictive power of the yield spread reaches a peak at a forecast horizon of around four to six quarters, which is fairly compatible with the traditional point of view on the length of the lags between monetary policy measures and real economic activity (Cozier and Tkacz (1994));
- among the components of aggregate demand, the yield spread forecasts consumption particularly well, and within this durable goods. Its link with investment is fairly weak over a one-year forecast horizon, but strengthens when the horizon is extended. These results are also compatible with the traditional transmission model, according to which monetary policy affects first consumption spending and then, via the accelerator effect, investment. In addition, the effect of the yield spread on the consumption of non-durables is rather low, which does not appear to be consistent with forecasts obtained using the CAPM (Cozier and Tkacz (1994));
- historically, the yield spread has proved on average to be a better advance indicator of output than of inflation, which suggests that the movements in the yield spread capture changes in real rates better than those in inflation expectations; moreover, for predicting inflation, the yield spread's maximum explanatory power is reached at a longer forecast horizon than for predicting real output. This is consistent with the chain of causality which runs from the yield spread to real economic activity and, finally, to inflation (Cozier and Tkacz (1994) and Day and Lange (1997)).

Furthermore, we have some results, which allow us to conclude that the yield spread may have been a better indicator of monetary policy than the traditional direct measures of the real interest rate:

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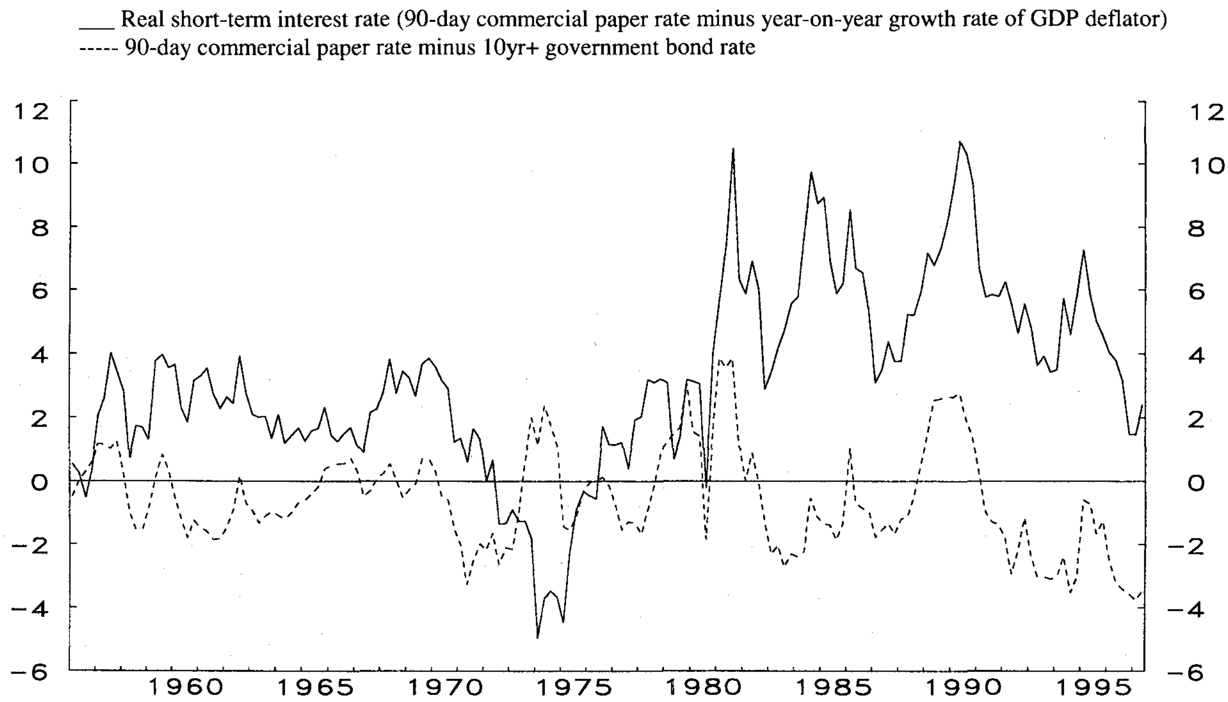
2 Certain observers sometimes wrongly think that the monetary conditions index (MCI) is used by the Bank as an indicator of monetary policy. In fact, the MCI is used as an operating target, the equivalent of the short-term interest rate at other central banks. As an indicator of monetary policy, the MCI suffers from the same shortcomings as the level of interest rates, since the values for the interest rate and the exchange rate which ensure a balance between supply and demand are not constant over time.

- historically, Canada's real short-term interest rate deviates from its average long-term value for prolonged periods (Graph 2), whereas the yield spread tends to return to its average value, as might be expected if it reflects the temporary effects of monetary factors (Clinton (1994-95));[3]
- the yield spread appears to partly solve the problem related to the use of traditional measures based on retrospective inflation expectations. The best example is that of the mid-1970s, where it seems improbable that ex ante real rates were so strongly negative;
- the yield spread gives better results than the real interest rate in an aggregated household consumption model, and provides good results for identifying monetary shocks in VAR models (Macklem (1995b) and Macklem, Paquet and Phaneuf (1996)).

Graph 2

Real short-term interest rate and spread between short and long rates



Based on these observations, the yield spread was chosen as the key monetary variable in QPM.[4] As well as isolating the monetary component of the changes in interest rates, the yield spread can serve as a guide with regard to the vigour with which short rates need to be raised in order to control the inflationary consequences of a shock. For example, a positive demand shock will provoke a rise in expected future real interest rates, and possibly in expected inflation. The extent to which long rates increase can thus provide information on the markets' perception of the authorities' determination to control inflation shocks.

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3 It should be noted, however, that the formal tests do not allow us to reject that the yield spread is non-stationary (see Section 3).

4 In QPM, this variable is used to capture the impact of interest rates on "consumption", which includes household consumption as defined in the national accounts, residential construction spending and changes in stocks. Obviously, the exchange rate also plays an important role in the model. See Coletti et al. (1996) for a description of the model.

1.4 Implications for the transmission mechanism

The fact that the yield spread is statistically superior to other interest rate measures suggests that monetary policy has little effect on real long-term rates, which would appear to be confirmed by the estimation results of Clinton and Zelmer (1997). Using a small-scale VAR, they find that short-term interest rate shocks brought about by Canadian monetary policy have very weak, indeed insignificant, effects on Canadian long rates.[5] However, economists generally maintain that it is long rates that most affect spending by households, given the structure of their balance sheets.[6] How, therefore, can one reconcile the above estimation results, which accord a major role to monetary policy in the explanation of economic fluctuations, with the negligible effect of monetary policy on long interest rates?

Table 2

VAR model results

August 1972 - December 1996

Terms for RL	**Maximum response of RL to a 100 b.p. innovation of R90**	**Probability Σ coefficient (R90 - R90US) = 0**
1 year	0.56	0.00
2 years	0.42	0.00
3 years	0.32	0.00
5 years	0.24	0.00
10 years	0.20	0.00
Long term	0.08	0.19

Note: The model is based on 90-day (R90) and long-term (RL) interest rate differentials between Canada and the United States.

Two factors are important in this regard. First, monetary policy influences household spending through various channels, notably via intertemporal substitution effects. This effect is independent of the balance-sheet structure. For example, if real short rates rise in relation to expected future rates (the yield spread narrows), this tends to dampen the demand for credit and delay consumption spending. Secondly, even if monetary policy has no major influence on long-term rates, its impact on medium-term rates is nevertheless substantial. A study by Montplaisir (1996-97) shows that contracts at three and five years currently account for the majority of households' financial liabilities. The rates associated with these maturities are therefore most likely to influence that sector's liquidity constraints and, when re-estimating the Clinton and Zelmer model, a shock to Canadian short-term rates has a significant effect on the rates for three and five-year bonds (Table 2).[7] Monetary policy may therefore also exert considerable cash-flow effects.

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5 These results must not be interpreted as signifying the existence of a weak correlation between the Canadian short and long rates. The tests presented in Section 3 in fact reveal an extremely close correlation between these rates. What the VAR indicates is simply that the independent effects of Canadian monetary policy on Canadian long rates are very weak.

6 The same could be said of firms' spending. However, it has always been very difficult to estimate significant interest rate effects for firms.

7 It is interesting to note that the results of the VAR indicate that the effects of a monetary policy shock on implied one-year forward rates are greater than might have been expected if the expectations hypothesis of the term structure of

1.5 Unanswered questions

Although we selected the yield spread for the projection model, questions remain as to its predictive power. More research is needed to develop a more solid theory to explain the cause-and-effect relationships at work. Certain results in particular raise questions:

- the yield spread fails to predict the growth of a particular spending category as precisely as the growth of global spending. Since monetary policy directly affects households' expenditure on durable goods, one would expect the yield spread to be able to forecast this more precisely than aggregate spending (Cozier and Tkacz (1994));
- over the period 1972-90, the predictive power of the spread between medium-term rates is just as high as that of the spread between short and long rates in the indicator models for the growth of real GDP (Table 1). One might expect, however, that the latter spread would yield better results as an indicator of monetary policy, since the short rate is controlled more closely by the Bank, while the long rate should represent the equilibrium rate better.

In addition, over the last few years, the indicator models based on the yield curve have substantially overpredicted the growth of economic activity. Since other economic models have also overpredicted growth, it is too early to tell whether these errors are symptomatic of a break in the relationship between the yield spread and economic activity. Can the move to low inflation and lower interest rates worldwide have affected the predictive power of the yield spread? Does this power vary according to the monetary policy stance? It is interesting to note that, in research aimed at testing the asymmetry effects of monetary policy, the conclusions vary depending on whether or not the yield spread is used as a measure of monetary policy. Using the spread, the results are more favourable to the asymmetry hypothesis.[8] It has yet to be demonstrated whether these results really do reflect the asymmetry in the effects of monetary policy, or whether they instead capture the asymmetry in the determination of long rates.[9]

According to Clinton and Zelmer (1997), one possible explanation for the forecast errors of the yield spread is the increase in the risk premium on Canadian long-term securities during the 1990s. When they include a variable in the indicator model to capture this risk premium, it has the expected sign and is significant, although it should be noted that this addition improves the forecasts obtained only slightly. At this stage of our research, however, we are not able to say whether this results from the fact that the variations in the term premium are not very important in explaining the forecast errors, or whether it simply reflects the difficulty in measuring this term premium precisely. Work is currently under way to improve our understanding of the determinants of the term premium and we will return to this topic in Section 3.[10]

2. The indicator role of the term structure of interest rates

In this section, we briefly examine how the information content of the term structure of interest rates is used by the Bank, from both a strategic and a tactical point of view.

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interest rates had been confirmed. These results are compatible with Shiller's hypothesis of the overreaction of long-term rates, or they may indicate that the liquidity effect influences the term premia.

8 See, in particular, Macklem (1995a) and Macklem, Paquet and Phaneuf (1996).

9 Remolona, Dziwura and Pedreza (1995) maintain, for example, that long-term rates contain more information when agents expect a tightening of monetary policy than when they expect an easing. In the first case, the changes in short rates dominate those in long rates, whereas, in the second case, the movements in the term premium are predominant.

10 In the context of the projection, adjustments are made to the equilibrium values for the yield spread in order to capture the variations in the term premium. This allows us to cancel out their effects in the model.

2.1 Real economic activity and inflation

Since monetary policy in Canada is oriented towards the achievement of inflation targets, the staff's economic projection and the constant monitoring of economic developments using different indicator variables are of prime importance. At the Bank, projections are made using a formal model, QPM. In addition, we have for some years been using indicator models based on financial variables in order to obtain alternative short-term forecasts of the path of real GDP and inflation. However, since monetary policy measures are aimed at influencing the inflation rate approximately six to eight quarters ahead, recent work on the indicator models has attempted chiefly to develop models with a forecast horizon of this length, with the purpose of cross-checking the projections made using QPM and thereby rapidly detecting potential errors.

As noted in the previous section, the research at the Bank has allowed us to conclude that the yield spread has, in the past, been an excellent advance indicator of real output growth. Its predictive power remains high even over a horizon of eight quarters, which makes it a particularly attractive indicator.[11] Since the indicator models have systematically overpredicted during the last few years, they have not played a front-line role in the conduct of monetary policy. We are continuing our research in order to improve our understanding of the source of these errors.

We also examined the relative performance of different financial variables in probit models aimed at determining the probability of a recession. The interest in this type of model lies, among other things, in the fact that it avoids the problem of illusory precision associated with point estimates. It is also possible that the yield spread is better at forecasting major variations in output growth, such as recessions. In studies currently under way, Atta-Mensah and Tkacz (1997) conclude that, among the indicators examined, the yield spread (from bonds at ten years and above to those at 90 days) is the best for forecasting recessions over a horizon of one to five quarters. It outperforms, inter alia, various measures of the level of nominal and real interest rates as well as equity indices and the monetary aggregate M1. The results of this model for a forecast horizon of four quarters are shown in Graph 3. These results are fairly consistent with those obtained in the United States by Estrella and Mishkin (1995) and for the Group of Seven countries by Bernard and Gerlach (1996).

Since what ultimately interests the Bank is determining the size of inflationary pressures in advance, we also examined the relative efficiency of different financial variables in forecasting periods of overheating. If we arbitrarily define such periods as those in which the output gap exceeds 2%, very preliminary studies suggest that the equity index would be a better predictor of these periods than the yield spread, which appears to have a fairly low predictive power. It is, however, too early to draw firm conclusions from these results; in particular, we have to test their robustness by using other definitions of periods of overheating.

Finally, Day and Lange (1997) have evaluated the ability of the yield curve to forecast future changes in inflation in Canada. They conclude that the slope of the yield curve for maturities of one to five years is a relatively reliable indicator of the future path of inflation at these horizons and that it contains different information to other indicators, such as the broad money aggregate M2+ and the output gap.[12] Nevertheless, the authors stress that, in the short term, the yield curve can vary as a result of temporary changes in real interest rates or term premia. Only lasting changes in the yield curve will be associated with similar changes in future inflation. Their results show that the explanatory power of the yield curve has increased considerably since the mid-1980s, probably because there have been no major supply shocks during this period. However, forecasts have

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[11] It is important to emphasise that, even if the yield spread is included in the projection model, it can also be used in indicator models since the projection model is much more complex and the results partly reflect expert judgement.

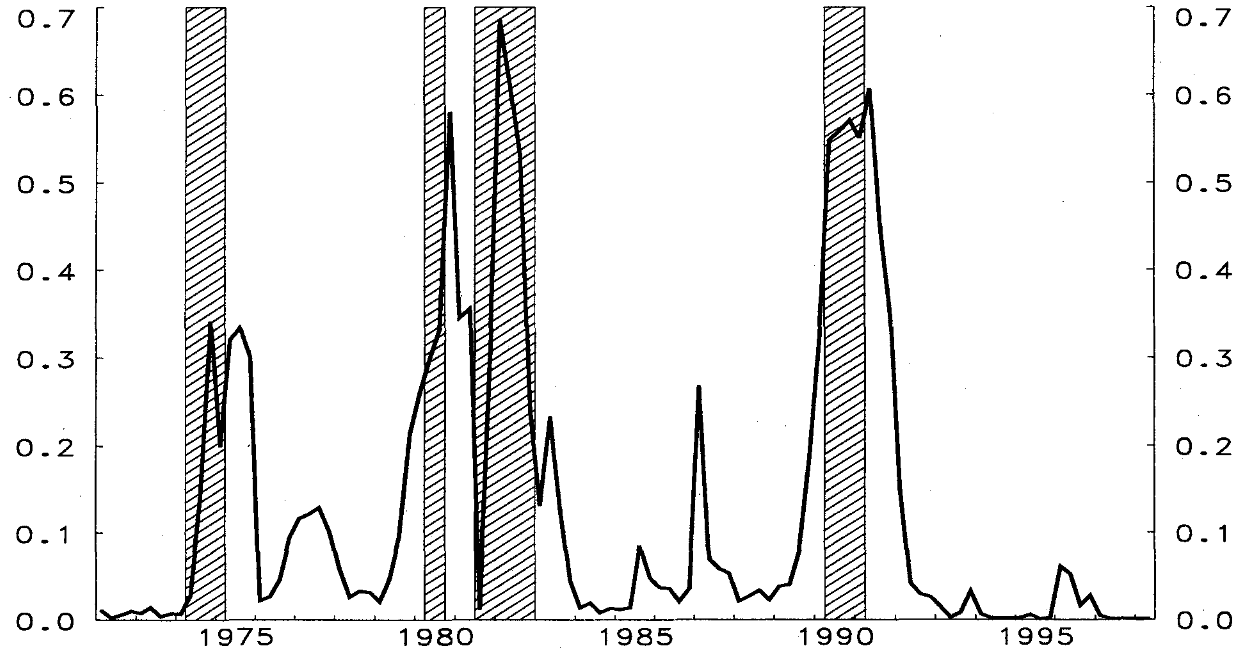
[12] The findings of Day and Lange are compatible with those of Mishkin (1990) for the United States and Gerlach (1995) for Germany. These last two authors conclude that the medium-term segment of the yield curve contains a great deal of information on future inflation.

deteriorated during the last few years, as can be seen from Graph 4. According to this model, inflation should accelerate by ½ to 1 percentage point from the end of 1998.

Graph 3

Four-quarter ahead probability of recession

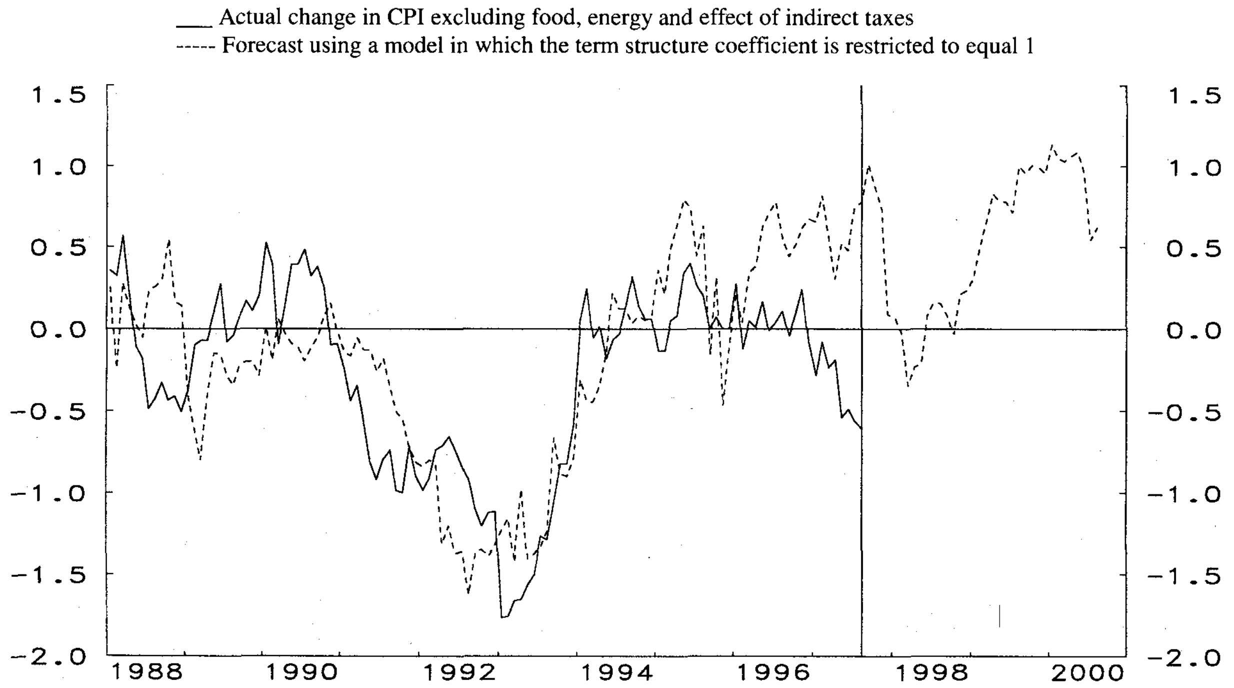
Out-of-sample forecast using 10yr+ bond rate less 90-day CP rate



Graph 4

Change in actual and predicted inflation

Three-year minus one-year inflation



2.2 The expected level of monetary conditions

At the Bank we also use the short interest rate segment of the term structure to measure the financial markets' expectations with regard to future three-month interest rates. The process consists of calculating the expected profile of three-month rates on the basis of the interest rates corresponding to forward interest rate agreements maturing in four, six, nine and 12 months, from which we subtract a representative term premium which varies according to the maturity of the agreement. The term premia are calculated using the average value of interest rate spreads over a long period that excludes certain episodes of high interest rate variability. The term premia also contain a variable (zero-centred) component obtained by estimating the cointegrating vectors linking the three-month rates with each of the forward rate agreements.

Together with the expected measure of the Canadian dollar exchange rate obtained from forward contracts, the measure of interest rate expectations is compared with that resulting from the Bank staff's economic projection, with the aim of evaluating the forecast level of monetary conditions relative to that expected by the financial markets. This comparison can be useful to the monetary authorities in their tactical decisions as to the appropriate moment to change the official interest rate.[13]

3. Tests of the term structure of interest rates

As we saw in Section 1, the different interpretations which can be placed on the role of the term structure of interest rates in the monetary policy transmission process are based in part on the expectations hypothesis (EH) of the term structure of interest rates. In its most general form, the EH states that each long-term interest rate represents the average of current and expected short-term interest rates over the life of the long-term security, plus a relatively stable term premium. In this section, we review the work recently undertaken to verify this hypothesis using Canadian data. We also examine the possibility that the presence of a risk premium on Canadian securities in international markets, as a result of Canada's high level of indebtedness, partly explains the variability of the term premium and the statistical rejection of the EH.

3.1 Tests of the expectations hypothesis of the term structure of interest rates

A number of tests were carried out to verify the EH using Canadian data. We divide these tests into three main groups, aimed at verifying: (i) whether long-term interest rates are unbiased predictors of future short-term interest rates; (ii) whether the long-term interest rate forecasts calculated using a model in which the EH is imposed permit adequate explanation of the long-term interest rates observed in the markets; and (iii) whether there is a long-term common trend between short and long interest rates. The third type of test does not constitute a direct verification of the EH, but the presence of a long-term common trend between the interest rates, which are non-stationary variables, is a necessary condition for the EH.

The type (i) tests of the EH start with estimating the following equation:

$$\frac{1}{k}\left[\sum_{k=0}^{k-1} i_{t+k}\right] - i_t = \alpha + \beta\left(i_t^k - i_t\right) + v_t \tag{5}$$

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13 For a discussion of tactical considerations, see Zelmer (1995).

To verify the EH, it is necessary to test the hypothesis $\beta = 1$; that is, that the spread between long and short rates is an unbiased predictor of the average of future short-term rates during the k periods to come, where k corresponds to the life of the long-term security.[14] We can also test the EH using the interest rates on forward agreements instead of long spot rates. In this case, the estimated equation takes the form:

$$i_{t+k} - i_t = \alpha + \beta\left(i(k)_t - i_t\right) + v_t \tag{6}$$

where $i(k)$ is the short-term interest rate on a forward rate agreement starting in k periods.

Table 3

HASTI tests in the short-term interest rate segment

	β	**Test: β = 1 (*p*-value)**	**Source; data**
One-day average interest rate forecast using changes in:			
1-month rate	0.86	0.26	Stréliski (1997); 1992:11:23-1996:10:07
2-month rate	1.03	0.83	Stréliski (1997); 1992:11:23-1996:10:07
3-month rate	1.02	0.94	Stréliski (1997); 1992:11:23-1996:10:07
One-month average interest rate forecast using changes in:			
3-month rate	0.86	0.33	Gerlach, Smets (1997); 1979:3:12-1996:7:15
6-month rate	0.72	0.19	Gerlach, Smets (1997); 1979:3:12-1996:7:15
12-month rate	0.62	0.01	Gerlach, Smets (1997); 1979:3:12-1996:7:15
Three-month interest rate forecast using changes in:			
3-month futures rate (maturing in 6 months)	0.73	0.48	This paper; 1990:01:01-1997:07:07
3-month futures rate (maturing in 9 months)	0.95	0.84	This paper; 1990:01:01-1997:04:07
3-month futures rate (maturing in 12 months)	1.10	0.68	This paper; 1990:01:01-1997:01:06

The results of the type (i) tests sometimes favour the EH, but they apply only to the short end of the term structure, with a maturity of less than 12 months. As Table 3 shows, the results of Stréliski (1997) indicate that one, two and three-month rates are unbiased predictors of the average call-money rate during the coming 30, 60 or 90 days, with the β coefficients linking the short and long rates varying between 0.86 and 1.02 and not statistically different from unity. However, as Stréliski points out, these coefficients are not stable over time. The results of Gerlach and Smets (1997) show that three-month interest rates are fairly good predictors of average one-month rates over the subsequent three months. However, the ability of the six and twelve-month rates to predict one-month rates is fairly imprecise, with the β coefficients deviating further and further from unity. Finally, our results indicate that the rates of future agreements are unbiased predictors of three-month interest

[14] It should be noted that the error term v_t is a moving average representation, the order of which depends on k and on the frequency of the data used. When testing the hypothesis $\beta = 1$, we must take this characteristic of the error term into account. The results presented use the Newey-West procedure, which corrects the variance-covariance matrix.

rates up to nine months in advance. However, like Stréliski we found that the β coefficients are unstable over time. For example, simply adding the year 1989 to our base sample reduces the value of the β coefficients by approximately 0.4 points. This instability is perhaps a reflection of the fragility of the EH.[15]

The type (ii) tests were developed by Campbell and Shiller (1987) and applied to Canadian data, inter alia, by Hardouvelis (1994) and Gerlach (1996). According to this methodology, one must first estimate a VAR with two variables, namely the change in short-term interest rates and the spread between the long and short interest rates. The VAR is then used to forecast future short-term interest rates, and these forecasts serve to calculate the *theoretical* values of the long-term interest rates under the EH with a constant term premium. Finally, statistical tests are used to compare the theoretical values for long-term interest rates with their observed values. Overall, the results of Hardouvelis (1994) and Gerlach (1996) tend to support the EH with Canadian data. However, the authors admit that the tests of the EH are not very powerful. Indeed, using an alternative hypothesis under this methodology, Sutton (1997) obtains results which tend if anything to reject the EH.[16]

The type (iii) tests for common trends are, in the first instance, aimed at testing the hypothesis of cointegration between short and long-term interest rates, and then, if there is cointegration, testing for the presence of a common relation [1, -1] between these rates.[17] In Table 4, we present the results of these tests applied to a large range of Canadian interest rate pairs, ranging from one and three-month rates to those at one, two, three, four, five and ten years.[18] We use monthly data covering the period 1972-96. The tests are obtained from estimating VECMs (vector error correction models) according to the methodology proposed by Johansen (1988) and Johansen and Juselius (1990). It should, however, be noted that the majority of the equations estimated in the VECMs, and shown in Table 4, suffer from a fairly severe ARCH-type error heteroskedasticity problem. This means that the cointegration tests and the hypothesis tests [1, -1] must be interpreted with caution.

The two tests used, *MV* and *Trace*, do not jointly support the presence of cointegration. Indeed, of the 28 interest rate pairs, there are only four for which both tests reject the absence of cointegration. Cointegration is present at the short end of the yield curve (between 30 and 90-day rates), but the hypothesis of a common relation [1, -1] in this segment is rejected.[19] Cointegration is also found in a section of the medium-term rate segment (between two, three and four-year rates), but there are no cointegration links between this segment and that of short rates, or with the long segment. As regards the long-term coefficients linking the different interest rate pairs, these are not too far from unity and lie between 0.79 and 0.99. In sum, fairly close relationships exist between short and long-term interest rates, but the absence of cointegration suggests that the term premium is variable and non-stationary.

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[15] Some may maintain that this instability reflects that the sample periods are too short to produce reliable statistical results rather than the fragility of the base hypothesis.

[16] According to Sutton (1997), the existence of a significant excess correlation between long-term bond yields in different countries is an indication that the EH is rejected, at least in terms of the formulation proposed by Campbell and Shiller.

[17] With interest rates often being regarded as non-stationary, the cointegration hypothesis [1, - 1] signifies, on the one hand, that permanent shocks affecting short-term interest rates (and, consequently, expectations of future short rates) will be reflected in corresponding changes in long-term interest rates. On the other hand, if there is cointegration [1, - 1], permanent shocks to long rates (which reflect the changes in expectations) should ultimately be reflected in changes in future short rates. Thus, cointegration [1, - 1] is a necessary condition for the EH.

[18] Table 4 is taken from Tkacz (1997).

[19] Gravelle (1997) arrives at similar conclusions when examining the relationships between three-month rates and the rates on forward agreements maturing in four, six, nine and twelve months.

Table 4

Cointegration test between interest rate pairs

System	*MV*	*Trace*	Long-term vector	Test [1, –1] (p-value)
[90 days, 30 days]	25.81*	30.84*	[1, -0.979]	0.03
[1 year, 30 days]	11.02	14.38+	[1, -0.943]	0.59
[2 years, 30 days]	10.44	14.64+	[1, -0.868]	0.35
[3 years, 30 days]	10.16	14.17+	[1, -0.831]	0.31
[4 years, 30 days]	10.15	13.97+	[1, -0.813]	0.31
[5 years, 30 days]	10.28	14.25+	[1, -0.804]	0.32
[10 years, 30 days]	9.36	13.78+	[1, -0.792]	0.43
[1 year, 90 days]	9.74	13.07	[1, -0.996]	0.97
[2 years, 90 days]	9.23	13.54+	[1, -0.919]	0.62
[3 years, 90 days]	9.38	13.57+	[1, -0.880]	0.52
[4 years, 90 days]	9.69	13.74+	[1, -0.862]	0.49
[5 years, 90 days]	9.91	14.12+	[1, -0.850]	0.48
[10 years, 90 days]	9.12	13.73+	[1, -0.836]	0.56
[2 years, 1 year]	9.57	13.43+	[1, -0.904]	0.05
[3 years, 1 year]	7.71	12.02	[1, -0.865]	0.16
[4 years, 1 year]	8.54	12.99	[1, -0.865]	0.24
[5 years, 1 year]	9.78	14.44+	[1, -0.860]	0.25
[10 years, 1 year]	10.29	15.07+	[1, -0.787]	0.16
[3 years, 2 years]	11.67	15.83*	[1, -0.958]	0.16
[4 years, 2 years]	13.19+	17.04*	[1, -0.939]	0.20
[5 years, 2 years]	12.37+	16.46*	[1, -0.933]	0.30
[10 years, 2 years]	10.53	15.91*	[1, -0.867]	0.28
[4 years, 3 years]	12.68+	16.26*	[1, -0.976]	0.28
[5 years, 3 years]	11.31	15.39+	[1, -0.968]	0.45
[10 years, 3 years]	9.56	14.82+	[1, -0.892]	0.30
[5 years, 4 years]	10.99	15.41*	[1, -0.988]	0.55
[10 years, 4 years]	9.45	14.20+	[1, -0.899]	0.19
[10 years, 5 years]	11.49	16.05*	[1, -0.928]	0.15

Notes: The systems are estimated on monthly data for the period 1972-96. The order of the system, or the p-value, equals 12. The statistics for *MV* and *Trace* allow testing for cointegration, using the null hypothesis of no cointegration (for more details see the notes to Table 5).

Among the reasons for the difficulties in accepting the EH (or in identifying the presence of a common trend between short and long rates), the possibility that the term premium is time-varying is the one which has aroused the most interest at the Bank recently. We have a large number of results that tend to support the hypothesis of a variable term premium, but little evidence of the underlying factors that might influence it. Thus, there are a number of results which show that the term premium depends on the conditional variability of interest rates themselves, which can be characterised by the distribution function of the shocks affecting the interest rates and by the rate determination process. In addition, as demonstrated by Lee (1995), in a general equilibrium model of the structure of interest rates with a monetary constraint, the term premium may depend on the conditional variability of the growth of economic activity and on the growth of the money supply. The empirical findings by Lee on the basis of US data appear to support his theoretical model. However, in a similar study Hejazi and Lai (1996) were not able to identify such a relationship using Canadian data. According to their results, the term premium is linked to the conditional variability of both interest rates themselves and the exchange rate. This last result surely deserves particular attention in future research.

When the premium is variable, rejection of the EH can also depend on the difficulty of forecasting future interest rate changes. In this regard, the work of Gerlach and Smets (1997) has shown that the tests tend not to reject the base hypothesis in countries where short-term interest rates are more easily forecastable, particularly those using a fixed exchange rate regime, which is, of course, not the case for Canada and the United States. According to their results, Canada is the country in which short-term interest rates are the most difficult to predict after the United States. Furthermore, it is well known that the EH is frequently rejected for the United States.

By the same token, it is possible that the difficulty of forecasting the future stance of monetary policy explains the rejection of the EH. The Bank of Canada has made considerable efforts in recent years to make monetary policy objectives and actions more transparent to the public.[20] In addition, research is currently under way to verify the potential effects of transparency on the forecastability of short and long-term interest rates in Canada.

To sum up, even if the EH is often rejected by Canadian data on the basis of statistical criteria, the fact remains that it constitutes an important economic hypothesis for explaining the path of long-term interest rates. Thus, the base model is not so much faulty as incomplete. In itself, the rejection of the EH does not, perhaps, pose a very serious problem for the monetary authorities. What presents a greater challenge, however, is understanding the reasons for the rejection of the base hypothesis. If the determining process of long-term interest rates changes over time, this would explain the rejection of the EH, while at the same time complicating the analyses of the monetary policy transmission process as well as the information content of the term structure of interest rates. This is why research is still in progress on testing the EH and understanding the reasons for its rejection. In the next section, we examine one potential cause.

3.2 The risk premium linked to public sector debt, the term premium and the common trend between short and long rates

Along with the difficulty of establishing stable links between short and long-term interest rates in Canada, various studies have shown the very strong substitutability between Canadian and US bonds, and the very close links which exist between the interest rates of these two countries.[21] More recently, studies have revealed that a *variable* risk premium could be attached to long-term Canadian bonds on the international markets, possibly because of the rise in Canada's public and external debt during the second half of the 1980s and the first half of the 1990s.[22] This risk premium, which we will

[20] See the paper by Clinton and Zelmer (1997) on this subject.

[21] See, inter alia, the studies carried out at the Bank of Canada by Caramazza et al. (1986) and Murray and Khemani (1989).

[22] See the studies by Orr et al. (1995) and Fillion (1996).

refer to here as the debt premium, arises in part out of the uncertainty that often attaches to the value of the currency of a heavily indebted country. In the light of these results, it would seem of interest to examine the hypothesis that the debt premium also influences the term premium. Tests of the EH (or tests for a common trend) which do not take account of this situation might tend to reject, spuriously, the base model.

The hypothesis we wish to examine derives from the following three long-run relations:[23]

$$i^k = if^k + \Delta^e z(l) + \Omega_l \tag{7}$$

$$i = if + \Delta^e z(s) + \Omega_s \tag{8}$$

$$if^k = if + \varphi \tag{9}$$

Equation (7) represents the hypothesis of uncovered interest parity between long-term rates in Canada and the United States. It states simply that long-term interest rates in Canada (i^k) are equal to long-term interest rates in the United States (if^k), plus the expected changes in the exchange rate over the life of long-term bonds ($\Delta^e z(l)$), plus a risk premium (Ω_l) which may correspond to the debt premium. Equation (8) represents the hypothesis of uncovered interest parity between short-term interest rates in Canada and the United States (i and if). It takes a similar form to equation (7). Equation (9) states that there is a long-run unit root between long and short-term interest rates in the United States, plus a term premium φ. Thus, the model is based on the hypothesis of a common trend between short and long rates in the United States, a hypothesis that is not necessarily accepted unanimously.[24]

By substituting equation (9) into (7) and subtracting (8), we obtain the following equation:

$$i^k - i = \varphi + \left[\Delta^e z(l) - \Delta^e z(s)\right] + \left[\Omega_l - \Omega_s\right] \tag{10}$$

Thus, we find in equation (10) that short and long-term interest rates have a unit root. We note that the spread between long and short rates can depend on a number of factors. For simplicity's sake, we assume that the expected short and long-term exchange rate changes are equal, in other words, $\Delta^e z(l) = \Delta^e z(s)$.[25] In addition, we assume that the term premium in the yield curve for the United States (φ) is predetermined and stationary, as is the risk premium incorporated in short-term interest rates in Canada (Ω_s).[26] These hypotheses allow us to obtain a stationary component for the term premium, which we define as $\Phi = \varphi - \Omega_s$. Finally, we assume that the risk premium in long-term

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23 The approach which follows is an approximation. It serves to illustrate the hypothesis we wish to examine rather than to obtain an exact formulation of the term structure. The time indices have been omitted in order to facilitate the notation.

24 Engsted and Tanggaard (1994) do not reject the hypothesis of a common trend between short and long-term interest rates in the United States on the basis of cointegration tests obtained from the estimation of VECMs. However, Gerlach (1996) and Gerlach and Smets (1997) strongly reject the EH using type (i) and (ii) tests applied to US data. The fact that the EH does not hold in the United States may well explain why it is not accepted in Canada either, given the strong substitutability between Canadian and US bonds, but we do not address this possibility in this paper.

25 This hypothesis may seem fairly extreme, but it is probably correct given the difficulty in finding an appropriate exchange rate forecast model.

26 We have certain reasons for believing that the spread between Canadian and US short-term interest rates is not stationary, and that this spread is linked to Canada's public and external debt ratios. However, this relationship appears to be unstable. This is why we favour the hypothesis that the risk premium incorporated in short-term bonds is stationary.

interest rates in Canada is variable and, more particularly, that it depends on Canada's public sector debt ratio $\Omega_l(D)$, where D is the ratio of public sector debt to nominal GDP.

Under these hypotheses, we obtain the following formulation for the term structure of Canadian interest rates:

$$i^k = i + \Phi + \Omega_l(D)) \tag{11}$$

To test this formulation, we use a set of cointegration and hypothesis tests obtained from the estimation of VECMs similar to those previously discussed. We apply these VECMs to interest rate measures from which we have previously subtracted the inflation rate for the preceding year, because this transformation reduces the problem of error heteroskedasticity we described earlier. The main estimation results are presented in Table 5. The estimation period runs from the first quarter of 1972 to the last quarter of 1994.

Table 5

Cointegration tests between interest rates (data adjusted for the previous year inflation rate)

System	**Cointegration tests**			**Univariate specification tests**		**Hypothesis tests**	
	H0	***MV***	***Trace***	**LB(24)**	**ARCH (5)**	**Non-constrained/constrained long-term vector coefficients**	***p*-value**
Tests of common trend							
(1) *rlcd, rscd*	$r=0$ $r\leq1$	9.81 3.01+	12.82 3.01+	0.51 0.87	0.21 0.27	1a. [1, -1.18]/[1, **-1**]	0.32
(2) *rlcd, rscd, ngl*	$r=0$ $r\leq1$	19.77+ 8.44	28.49+ 8.71	0.34 0.44 0.30	0.08 0.04 0.46	2a. [1, -0.69, -0.03]/[1, **-1,** 0.01] 2b. [1, -1, -0.01]/[1, -1, **0**] 2c. [1, -0.69, -0.03]/[1, -0.75, **0**]	0.03 0.78 0.18
(3) *rlcd, rscd, prime*	$r=0$ $r\leq1$	18.83+ 9.51	30.30+ 11.47	0.66 0.99 0.31	0.13 0.46 0.62	3a. [1, -0.69, -1.80]/[1, **-1,** 0.88] 3b. [1, -0.69, -1.80]/[1, -0.87, **1**] 3c. [1, -0.87, -1]/[1, **-1**, -1]	0.04 0.11 0.20
Tests of non-covered interest parity hypothesis (see Annex)							
(4) *rlcd, rleu*	$r=0$ $r\leq1$	6.68 1.53	8.22 1.53	0.63 0.82	0.06 0.90	4a. [1, -1.17]/[1, **-1**]	0.40
(5) *rlcd, rleu, ngl, nfl*	$r=0$ $r\leq1$	31.20* 12.43	53.59* 22.39	0.42 0.63 0.05 0.47	0.19 0.45 0.33 0.91	5a. [1, -1.02, 0.01, -0.17]/ [1, **-1,** 0.01, -0.17] 5b. [1, -1, 0.01, -0.17]/ [1, -1, **0**, -0.15] 5c. [1, -1, 0.01, -0.17]/ [1, -1, -0.03, **0**]	0.87 0.76 0.03

Notes: The systems are estimated on data from 1972Q1–1994Q4. The order of the system, or the *p*-value, equals 5. All equations are estimated including a constant term. The statistics for *MV* and *Trace* allow testing for cointegration, with H0: $r = 0$ indicating that we are testing the null of no cointegrated vectors. If the null cannot be rejected, i.e. there is at least one vector of cointegration, a step-wise procedure is used to verify that there is no more than one vector. Thus, if H0: $r \leq 1$ cannot be rejected and H0: $r = 0$ has already been rejected, there is at most one cointegrating vector. On the other hand, if H0: $r \leq 1$ is rejected, there is more than one vector. "+" indicates statistical significance at a confidence level of more than 90% while "*" indicates a confidence level above 95%, with critical values taken from Osterwald and Lenum (1992). The statistics LB(24) and ARCH(5) test for, respectively, autocorrelation and heteroskedasticity in the error terms, using a Chi-square test. The first hypothesis tested is that there is a unitary relationship between pairs of interest rates (indicated by 1, **-1**, d1), with bold figures referring to the hypotheses being tested. We also test the hypothesis that other variables may have a significant influence different from 0 or 1 (indicated by (1, -1, **0**) or 1, -1, **-1**)).

We first use system (1) to examine the simple relation between long-term (i^k) and short-term (i) interest rates in Canada. The results of the *MV* and *Trace* tests do not allow the absence of cointegration to be rejected. As with the results presented in the previous section, this indicates that the short and long-term interest rates are not cointegrated, even though there appears to be a fairly close relation between i^k and i.

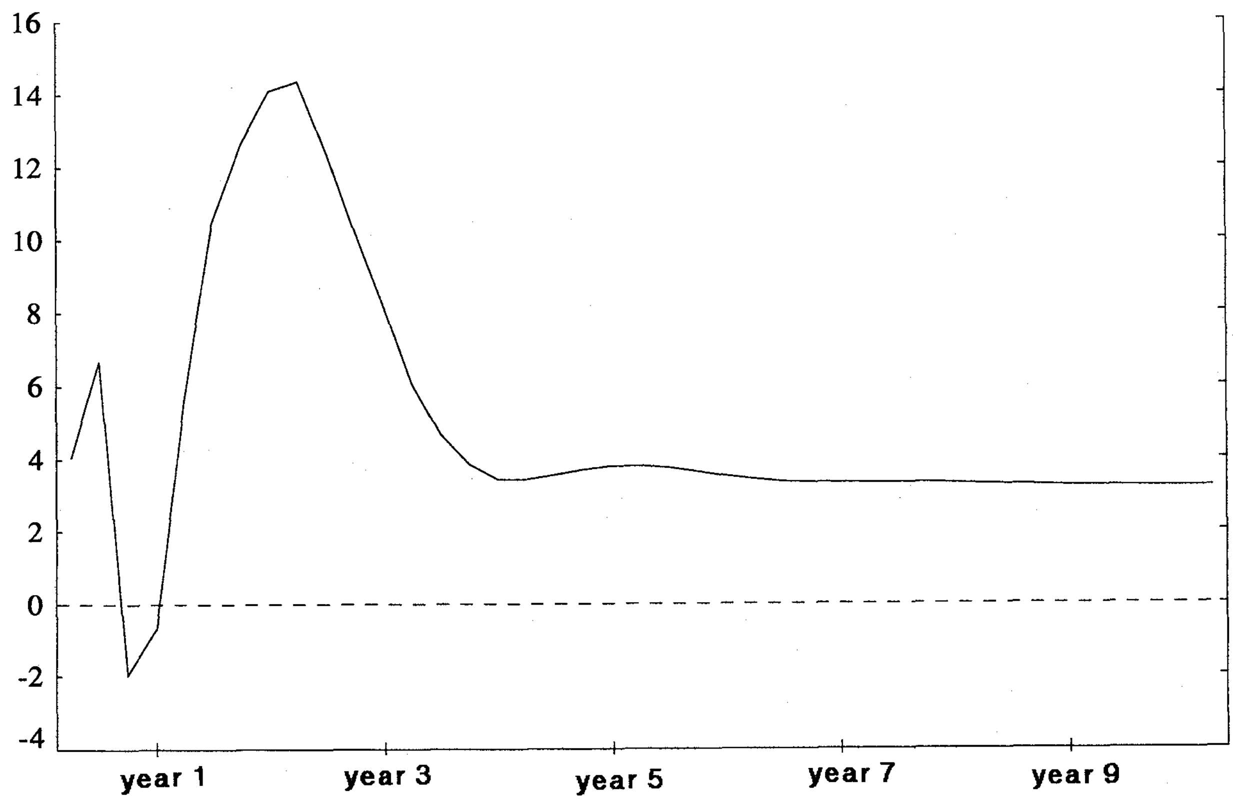
The results of system (2) show a cointegrating relation between i^k, i and the public sector debt ratio *ngl*. However, we can easily reject the hypothesis of a unit root between i^k and i in this model (system (2a)). In addition, in the system in which the unit root between i^k and i is imposed, the public sector debt ratio has no significant effect (system (2b)). Although the public sector debt effect is also insignificant in the system in which the unit root between i^k and i is not imposed (system (2c)), it is not negligible economically, since each 1 percentage point increase in the public sector debt ratio causes real interest rates to rise by 3 basis points in the long term (Graph 5).

These results suggest that there is a non-stationary component of the term premium which is linked to the public sector debt ratio. However, as we have just seen, the effect of this component is difficult to measure precisely. By the same token, the work recently carried out by Fillion (1996) suggests that, in order to evaluate the effects of the public sector debt ratio on the risk premium incorporated in Canadian long-term interest rates, account must be taken of: (a) the effects of the public sector debt on Canada's external indebtedness; and (b) the close cointegrating relation which exists between long-term interest rates and external debt in Canada.

Graph 5

Effect of a 1% increase in the public sector debt/GDP ratio on long-term interest rates

In basis points



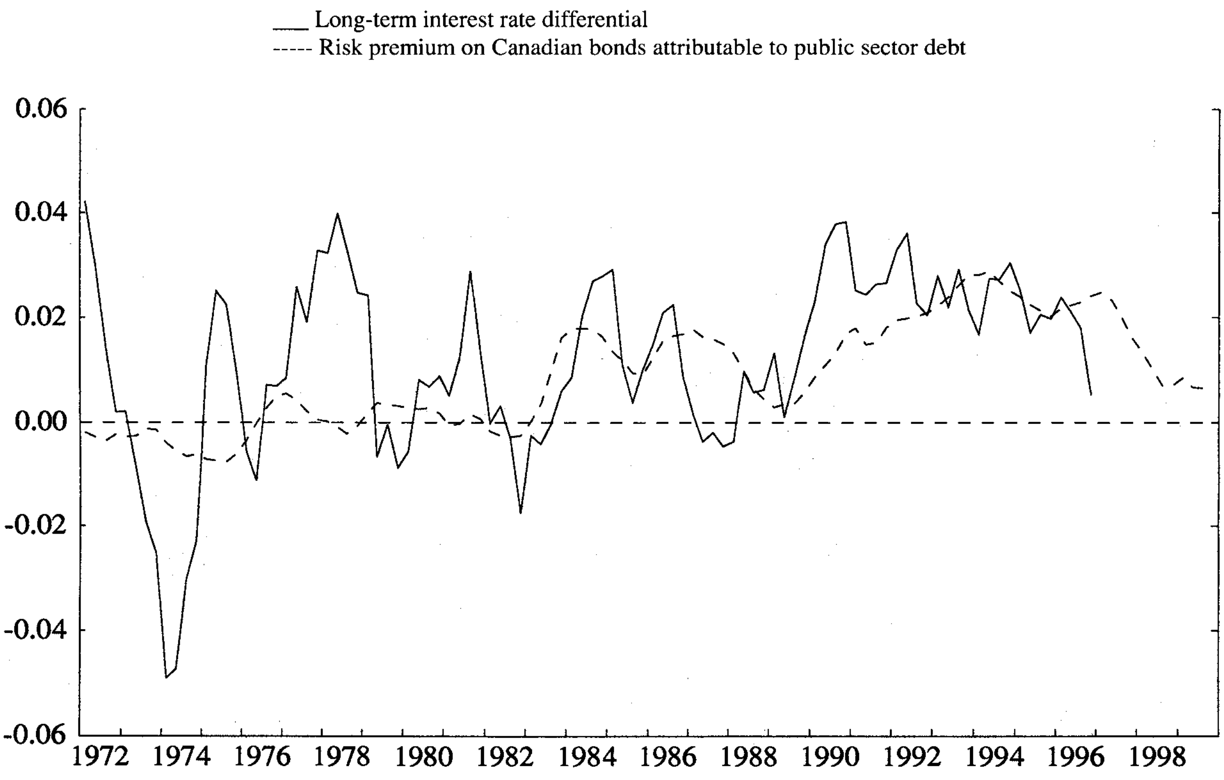
We simulated the VECM estimated by Fillion in order to evaluate the effect of the change in the public sector debt ratio on the risk premium since the beginning of the 1970s, after which we introduced this measure of the debt premium into the relation for the term structure of

Canadian interest rates.[27] The results of this simulation are shown in Graph 6. (We assume that the debt ratio during the period 1997-99 will decrease at the same pace as it increased during the three years prior to its 1996 peak.) We note first of all that the risk premium attributable to the public sector debt coincides fairly well with the major changes in the spread between Canadian and US interest rates, in particular since the beginning of the 1980s. For example, the spread between Canadian and US rates narrowed by approximately 150 points during the second half of 1996. However, according to our simulations, the fall in the debt premium ought to have been observed in 1997, when the public sector debt ratio effectively started to decrease.

In the last stage of this study, we introduce the measure of the debt premium (*premium*) presented in Graph 6 into the relation for the term structure of Canadian interest rates. The results (Table 5 (system (3)) indicate the presence of a cointegrating relation. We note, however, that the unit root between short and long-term interest rates is rejected (system (3a)).[28] When we modify the sequence of hypothesis tests, we observe that we cannot reject the hypothesis that the debt premium has a one-for-one effect on long-term interest rates (system (3b)), nor can we reject the unit root between short and long-term interest rates (system (3c)).

Graph 6

Long-term interest rate differential between Canada and the United States (adjusted for inflation differential) and estimated risk premium on Canadian bonds



This last set of results shows that a common trend exists between short and long-term interest rates when we take into account the presence, in the long rate segment, of a variable risk

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[27] For interested readers, a brief overview of Fillion's (1996) results are provided at Annex 1.

[28] In the system in which the unit root between short and long rates is imposed, the variable *premium* shows an effect significantly different from zero, but not different from one (results not reported here).

premium linked to Canada's public sector debt ratio. This suggests that the existence of the risk premium might also help explain the rejection of the EH on the basis of the usual tests.

Conclusions

The aim of this paper has been to put into perspective the empirical results obtained at the Bank of Canada and elsewhere on the subject of the information content of the term structure of interest rates and to describe how this information is currently used in the conduct of monetary policy in Canada. From the wealth of financial instruments whose prices may contain useful information for a central bank, we have confined ourselves to examining the term structure of interest rates because this is currently the most important source of information for the Bank of Canada and has been the subject of a number of studies using Canadian data.

A large amount of research is currently being carried out at the Bank aimed at extracting information from the prices of other financial assets. This research, and the relevant research undertaken elsewhere in Canada, will be presented in May 1998 at a conference organised by the Bank. Among the questions which need to be examined is that of the information on the distribution of probabilities relating to exchange rate expectations that can be extracted from the prices of option contracts, plus the information on inflation expectations that can be extracted from long-term interest rates.

Annex: Canadian debt and its effects on long-term interest rates

This annex provides an overview of the results in Fillion (1996) which were used in Section 3 of this paper to calculate a measure of the risk premium on Canadian bonds linked to the development of the public sector debt ratio.

The results from system (4), presented in the second half of Table 5, show that we cannot reject the absence of cointegration between real long-term interest rates in Canada i^k and the United States if^k. System (5), on the other hand, indicates a close cointegrating relation between i^k, if^k, the public sector debt ratio *ngl* and the Canadian external debt ratio *nfl*. This result in favour of cointegration is obtained because the system contains two important endogenous variables, i^k and *nfl*.[29] Indeed, it would appear crucial to take account of the endogenous character of external debt in order to identify a cointegrating relation between interest rates in Canada and the debt variables used to approximate the risk on Canadian bonds. Furthermore, the results show that if^k and *ngl* are exogenous variables, in the weak sense, in this system.[30] The results also reveal a very close relationship between Canadian and US long-term interest rates (system (5a)). In addition, they indicate that *ngl* has no significant effect (system (5b)), whereas *nfl* is significant at a confidence level of over 95% (system (5c)).

29 The endogenous or exogenous character is identified using tests of significance of the adjustment parameters which are attached to the cointegrating vector in each equation.

30 While the public sector debt may be exogenous in the weak sense, it is a little difficult to believe that it is strictly exogenous, if only because of the effect that interest rate changes can have on the servicing of public sector debt and, consequently, on deficits and indebtedness. This relation is found in our systems of equations, but these systems are not well adapted to examining this particular question. See Fillion (1996) for more details on this subject.

It would therefore appear that a cointegrating relation exists between Canada's external debt ratio and Canadian interest rates. However, since the external debt variable is endogenous, it is difficult to quantify the role of this variable in the measure of the risk premium. On the other hand, the public sector debt ratio, which is more exogenous, does not seem to have a significant direct effect on the risk premium. Nevertheless, this ratio may have a major effect on the risk premium to the extent that it influences the development of the external debt ratio. In order to evaluate the effect of public sector debt on the risk premium for Canadian bonds, we estimated the VECM including the variables i^k, if^k, *ngl* and *nfl* by postulating that *ngl* and if^k are exogenous, and we simulated it for the values observed for the public sector debt ratio since the beginning of the 1970s.[31] The results of this simulation are presented in Graph 6. We discuss them in more detail in the paper.

Another way of evaluating the effect of the public sector debt ratio on the risk premium is to submit the system of equations to a representative shock of 1 percentage point of *ngl* (see Graph 5 of this paper).[32] In this system, each 1 percentage point rise in *ngl* has the long-run effect of increasing the external debt ratio by 0.22 percentage points. On the basis of the estimated cointegrating relation, it is easy to establish that the 1-point shock to the public sector debt ratio, given its effect on *nfl*, causes an increase in the risk premium for Canadian long-term bonds of 3.1 basis points after a certain time has elapsed. The dynamic profile of this effect is shown in Graph 5. Although the impact on the risk premium is imprecise during the first year, it is particularly high during the second and third years. After the third year, the simulations converge rapidly towards the long-run value. The strong rise in the risk premium during the intermediate period may reflect the reaction of financial market participants to the uncertainty surrounding the links between a rise observed in the debt ratio and its expected future path.

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31 The estimated system thus contains only two equations, that of the first difference of *rlcd* and that of the first difference of *nfl*. Both estimated equations contain the first four lags of each of the four system variables, as well as the contemporaneous change in the two exogenous variables. In addition, both equations contain the long-term vector linking the levels of *rlcd*, *rleu*, *ngl* and *nfl*. In this vector, the unit root between *rlcd* and *rleu* is imposed during the estimation.

32 In order to create a representative shock, we estimated an equation Δngl, the formulation of which is similar to that of the two other equations of the VECM, and submitted it to a 1% shock. Thus, this shock takes account of the dynamics of the equation Δ *ngl*, as well as the presence of lags of Δ *rlcd* and Δ *nfl* in this equation.

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What do asset price movements in Germany tell monetary policy makers?

Dietrich Domanski and Manfred Kremer*

Introduction

Asset prices can play a twofold role in monetary policy. First, they may be seen as important elements in the chain along which monetary policy stimuli are transmitted to the real economy. From this perspective, asset price movements cause changes in aggregate demand or the price level through substitution, income and wealth effects. If these structural relationships were stable and could be estimated reliably, asset prices could be used as indicators of, or even target variables for, monetary policy. Second, they may be seen as predictors of the future course of the economy, independently of their active role in the transmission process. This view does not depend on the causal influence of asset prices on the macroeconomic variables to be predicted. Instead, it takes due account of the fact that the price of rationally valued assets should reflect the expected path of the asset's income components and the equilibrium returns used for discounting the future stream of income. If these expectations were influenced by the anticipated development of certain macroeconomic fundamental factors, and if, furthermore, market expectations were not systematically biased, asset prices could be used by the central bank as predictors of real activity and inflation.

The monetary policy implications of both roles depend crucially on the informational efficiency of asset markets. Market inefficiencies would cause asset prices to deviate from their fundamental values, distorting their informational content and their indicator quality. Furthermore, if asset prices play an important role in the transmission process, mispricing may adversely affect economic activity and price stability. The main body of this paper is devoted to assessing the predictive power or the informational content, respectively, of dividend yields and the term structure spread to draw some preliminary conclusions about the efficiency of the stock and government bond markets in Germany.

The theoretical framework is provided by the rational valuation approach. Applied to the bond market and the stock market, this approach leads to the expectations hypothesis and the dividend discount model, respectively, both on the assumption of rational expectations. The informational content is judged by metrics from univariate regression techniques using short and long-horizon measures for future inflation, stock returns, dividend growth, and interest rate changes as dependent variables and the spread or the dividend yield as regressors. The paper closes with some implications of the results for monetary policy.

1. Pricing stocks and bonds with the rational valuation approach

The value of financial assets generally depends on the future stream of payments the holder is entitled to receive. Hence, it is economically reasonable to calculate an asset's fundamental value as the discounted present value of the expected stream of income. The discount rate used can be interpreted as the required (expected) rate of return which attracts investors to hold the asset in their portfolios. In an informationally efficient market, an asset's actual market price should then equal its fundamental value as calculated by all or the marginal investor depending on whether expectations are assumed to be homogeneous or not. Thus, testing the informational efficiency of asset prices requires

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* The views expressed in the paper are those of the authors and not necessarily those of the Deutsche Bundesbank.

an assumption about the behaviour of equilibrium returns and a hypothesis as to how market agents form expectations.

1.1 Stock pricing

Applied to the stock market, this general valuation approach is the dividend discount model. We can derive it starting with the approximation formula for the continuously compounded one-period return h_{t+1} on stocks as suggested by Campbell and Shiller:[1]

$$h_{t+1} = k + \rho p_{t+1} + (1-\rho)d_{t+1} - p_t \tag{1.1}$$

with h_{t+1} = approximate continuously compounded (or logarithmic) one-period return on stocks over the holding period t+1; p_t = log of stock price measured at the end of period t; d_{t+1} = log of dividend paid out before the end of period t+1; $\rho \equiv 1/(1+\exp(\overline{d-p}))$, where $\overline{d-p}$ = average log dividend yield; and $k \equiv -\log(\rho)-(1-\rho)\log(1/\rho-1)$.

Equation (1.1) provides a loglinear relation between stock prices, returns and dividends, which is more convenient for calculation purposes if equilibrium returns are allowed to be time-varying. It is a first-order linear difference equation in the stock price. Solving forward and imposing the terminal condition $\lim_{j\to\infty} \rho^j p_{t+j} = 0$, yields:[2]

$$p_t = \frac{k}{1-\rho} + \sum_{j=0}^{\infty} \rho^j \left[(1-\rho)d_{t+1+j} - h_{t+1+j}\right] \tag{1.2}$$

Equation (1.2) is a mere identity, which says that today's stock price is high if future dividends are high and/or future returns are low. By applying the conditional expectations operator $E_t x_{t+1} = E[x_{t+1}|\Omega_t]$ (with Ω_t the market-wide information set available at the end of period t) and the law of iterated expectations, equation (1.2) can be changed to an ex ante relationship: [3]

$$p_t = \frac{k}{1-\rho} + \sum_{j=0}^{\infty} \rho^j \left[(1-\rho)E_t d_{t+1+j} - E_t h_{t+1+j}\right] \tag{1.3}$$

Further assuming homogeneous expectations on the part of all market participants and instantaneous market clearing, the log stock price always equals its single fundamental value, which in turn is the specifically weighted, infinite sum of expected log dividends discounted by principally time-varying expected equilibrium returns. Thus, equation (1.3) just represents the dividend discount model. Combined with rational expectations, it is also a valid representation of the "rational valuation formula" (RVF) for stocks.[4]

The loglinear approximation framework has two important advantages: first, it allows a linear and thus rather simple, analysis of the stock price behaviour. Second, it conforms with the empirically plausible assumption that dividends and stock returns follow loglinear stochastic

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1 See Campbell et al. (1997), pp. 260-2.

2 This terminal condition rules out rational bubbles that would cause the log stock price to grow exponentially forever at rate $1/\rho$ or faster (Campbell et al. (1997), pp. 262 f.).

3 In technical terms, the law of iterated expectations can be expressed as $E_t[E_{t+i-1}h_{t+i}] = E_t h_{t+i}$ which may be interpreted as a consistency condition under rational expectations.

4 See Cuthbertson (1996), who applies the RVF to various financial instruments (stocks, bonds, foreign exchange).

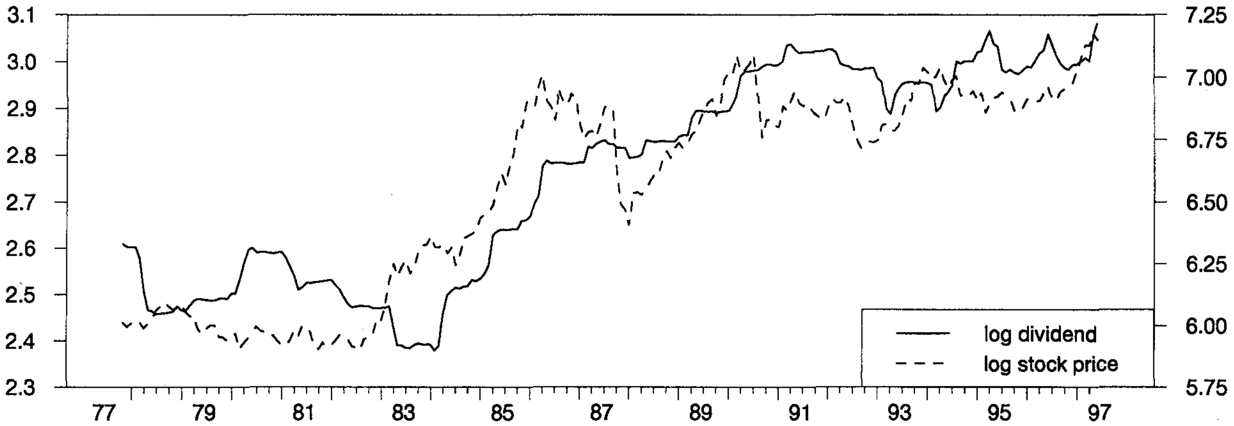
processes. For the empirical analysis it turns out to be advantageous to rearrange equation (1.3) such that the log dividend yield (or log dividend-price ratio) is singled out as the left-hand variable:

$$d_t - p_t = -\frac{k}{1-\rho} + \sum_{j=0}^{\infty} \rho^j (-E_t \Delta d_{t+1+j} + E_t h_{t+1+j}) \tag{1.4}$$

Figure 1

Dividends (left-hand scale) and stock prices (right-hand scale)

Nominal values (DM) in logs; November 1977 to June 1997



The empirical evidence generally suggests that the logs of dividends and stock prices follow non-stationary I(1) processes (see Figure 1). Dividend changes (the first differences) are therefore I(0) or stationary, as are the one-period stock returns. Thus, the right-hand side of equation (1.4) – a weighted sum of (expected) dividend changes and stock returns – should also be stationary. Dividends and stock prices must then cointegrate so that the (log) dividend yield can form a stationary process, too. If these stationarity assumptions were true, equation (1.4) would only consist of stationary variables and could be used for regression analysis without any further data transformations or use of non-standard distribution theory.

1.2 Bond pricing

Now we turn to the RVF for bonds. Since our analysis of the German bond market is based upon estimated spot rates (zero coupon rates), we start with the definition of the one-period return on a pure discount bond:

$$h_{t+1}^{(n)} = \ln(1 + H_{t+1}^{(n)}) = \ln P_{t+1}^{(n-1)} - \ln P_t^{(n)} \tag{1.5}$$

with $h_{t+1}^{(n)}$ = continuously compounded (or log) one-period return on a pure discount bond over the holding period t+1; $P_t^{(n)}$ = price of an n-period pure discount bond measured at the end of period t.

To cast equation (1.5) in terms of continuously compounded spot yields $z_t^{(n)}$, we substitute out bond prices by using the relation $\ln P_t^{(n)} = \ln M - n\ln(1 + Z_t^{(n)}) = \ln M - nz_t^{(n)}$. M is the redemption price of the n-period bond and $Z_t^{(n)}$ is the simple spot rate. Equation (1.5) then becomes:

$$h_{t+1}^{(n)} = nz_t^{(n)} - (n-1)z_{t+1}^{(n-1)} \tag{1.6}$$

The different theories of the term structure of interest rates are now based on different assumptions about the required or expected one-period return that attracts investors to hold an n-period bond over one period. We assume that investors require a rate of return which exceeds the one-period risk-free rate r_t by a term premium $T_t^{(n)}$:[5]

$$E_t h_{t+1}^{(n)} = E_t\left[n z_t^{(n)} - (n-1) z_{t+1}^{(n-1)}\right] = r_t + T_t^{(n)} \tag{1.7}$$

or $$n z_t^{(n)} = (n-1) E_t z_{t+1}^{(n-1)} + r_t + T_t^{(n)} \tag{1.8}$$

Now leading (1.8) one period, applying the law of iterated expectations and substituting the result into equation (1.8) gives:

$$n z_t^{(n)} = (n-2) E_t z_{t+2}^{(n-2)} + r_t + E_t r_{t+1} + T_t^{(n)} + E_t T_{t+1}^{(n-1)} \tag{1.9}$$

Further substituting and noting that $(n-j) E_t z_{t+j}^{(n-j)} = 0$ for $j = n$, we finally obtain a familiar term structure relationship which also represents the RVF for bonds:[6]

$$z_t^{(n)} = E_t\left[\frac{1}{n}\sum_{i=0}^{n-1} r_{t+i}\right] + E_t\left[\frac{1}{n}\sum_{i=0}^{n-1} T_{t+i}^{(n-i)}\right] = E_t\left[\frac{1}{n}\sum_{i=0}^{n-1} r_{t+i}\right] + E_t \phi_t^{(n)} \tag{1.10}$$

with $\phi_t^{(n)}$ = the average risk premium on the n-period bond until it matures. The n-period long-rate equals a weighted average of expected future short rates plus the expected average risk premium. But this equation is non-operational unless we assume a specific form of the term premium.[7] Different assumptions about the term premia also characterise the different term structure theories. For example, the pure expectations hypothesis (*PEH*) rests on the assumption of zero term premia for all maturities, while the expectations hypothesis (*EH*) only requires constant term premia which are the same for all maturities.[8]

Under empirically plausible assumptions about the time-series characteristics of interest rates, the following rearrangement of equation (1.10) leads to a stationary transformation, which is now widely used for regression purposes:[9]

$$S_t^{(n,1)} = z_t^{(n)} - r_t = \sum_{i=1}^{n-1} (1 - i/n) E_t \Delta r_{t+i} + E_t \phi_t^{(n)} \tag{1.11}$$

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5 The expected excess return may generally be called a risk premium. But since the yield data we use are for government bonds only which carry little or no default risk, the remaining risk of such bonds mainly arises from different *terms to maturity*. The expression "term premium" draws on this fact (see Cuthbertson (1996), p. 214).

6 The RVF for coupon-paying bonds is very similar to the formula for stocks. Uncertain dividend streams in the latter case are replaced by known coupon payments over a limited period of time, and, at maturity, the also known nominal value will be redeemed. This certain stream of (nominal) income has to be discounted using consecutive expected one-period returns required by the investors to hold the bond over its time to maturity, just as in the case of stocks. For pure discount bonds, only the redemption price has to be discounted to get the fundamental bond value and thus the RVF.

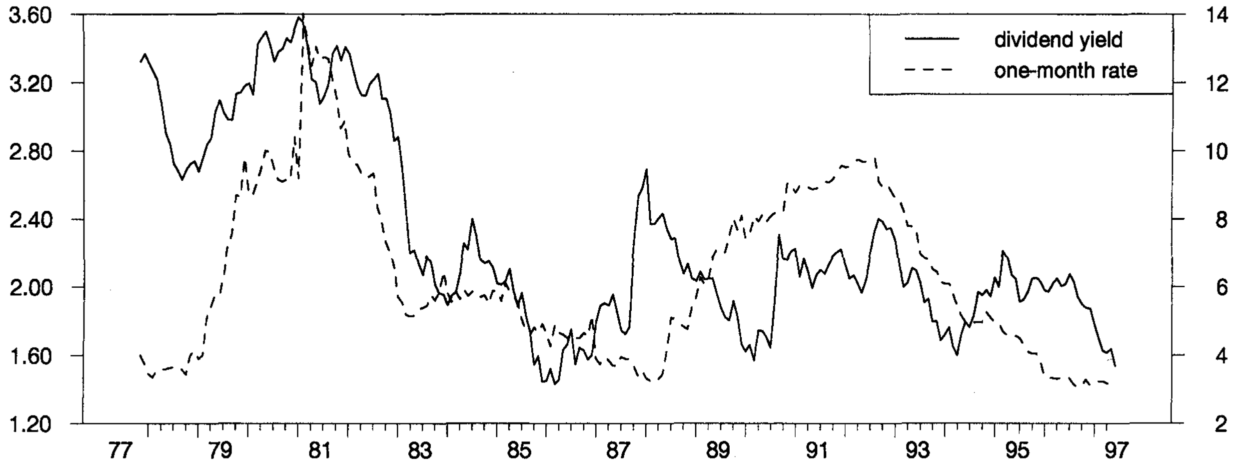
7 See Cuthbertson (1996), p. 225.

8 For a short survey of different term structure theories see, e.g., Cuthbertson (1996), pp. 218-23.

9 Although there are theoretically strong reasons for regarding interest rates as stationary variables, conventional integration tests most often suggest interest rates to be near-integrated variables whose time-series behavior may better be represented by non-stationary I(1) processes, at least in finite samples of typical size.

Figure 2

Dividend yield (left-hand scale) and one-month interest rate (right-hand scale)



Hence the spread between a long rate and a short rate should reflect the agents' expectations about future changes in the short rate and, under the expectations hypothesis, a constant term premium $\phi^{(n)}$ (see Figure 2). This is essentially an arbitrage condition saying that the investment in the long bond should earn the same return as successive short-term investments plus a risk premium that compensates for the capital risk incurred by holding the long bond.

2. Econometric evidence on the informational content and efficiency of German stock and bond market prices

The study of prices of long-term assets is intimately related to the study of long-horizon asset returns.[10] As equation (1.3) or, analogously, (1.4) shows, an infinite sum of future dividends enters into the calculation of the fundamental share value. Thus, the dividend of a single period can only be a small fraction of the stock price. Persistent changes in dividends therefore have a much larger influence on the stock price than do temporary dividend movements. A similar insight applies to changes in the discount rate used to value any financial asset.

This general conclusion provides the basis for the econometric analysis of this section. If dividend growth and discount rates follow predictable patterns, and if agents' expectations are not systematically biased, then the actual prices of longer-term assets like stocks and bonds should on average give useful information about the future course of asset returns or other variables correlated with the return process. It is intuitively plausible from the *RVF*s that in this case the forecast performance of current asset prices should generally be better for longer-term return measures (average returns), since these make up a larger part of the asset's calculated equilibrium price, and are, moreover, presumably less susceptible to large one-time shocks and peso effects than highly volatile short-term returns.[11]

In the following, long-horizon regressions are employed to determine the informational content of stock and bond market indicators regarding future stock returns, dividend growth, and short-term interest rate changes, respectively. Future ex post returns or short-rate changes measured over varying horizons are regressed on the current dividend yield or interest rate spread. The forecast

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[10] See Campbell et al. (1997), p. 253.

[11] See Kaul (1996), p. 284.

performance (predictive content) of each regression then helps to evaluate whether dividend yields or spreads better reflect *correctly anticipated* developments over longer or shorter horizons. This regression framework does not presuppose any specific model of the equilibrium return process. Thus, partial forecastability of returns (or excess returns) given lagged information, may indicate that conditional expected (equilibrium) returns are not constant, but vary over time, perhaps driven by time-variation in risk premia.[12] In addition, the predictive content of the same financial market indicators for future developments of macroeconomic variables like inflation or output generally provides some stylised facts about which fundamental factors are likely to determine equilibrium asset returns. In the present case, we ask about the informational content regarding inflation, since this is the most important variable from a monetary policy point of view.[13]

With respect to market efficiency, the long-horizon regressions for stock returns can be used to test the null hypothesis of constant equilibrium returns. Under this "traditional" hypothesis, future returns in excess of a constant should be unpredictable regardless of the return horizon and the information variables used.[14] In this single-equation setting, the unpredictability of stock returns can easily be tested by zero coefficient restrictions. However, in line with modern economic theory and the overall empirical evidence, it is now commonly believed that equilibrium returns vary over time. In this case, only returns in excess of the time-varying equilibrium component should be unpredictable. Efficiency tests under this assumption thus require a proxy for expected equilibrium returns. A short-term interest rate (the risk-free rate corresponding to the time-horizon over which returns are measured) is sometimes used for that purpose. As demonstrated above, this idea of constant equilibrium excess returns over a short-rate, applied to the bond market, leads to the expectation hypothesis of the term structure. Testing this hypothesis, which will be done below, is tantamount to testing bond market efficiency within the present framework.

Finally, a few comments on the data.[15] The RVF will not be applied to individual instruments but to broad portfolios of German stocks and bonds. While it is rather uncontroversial to refer to "average" bond yields calculated from a basket of homogeneous bonds (with comparable terms to maturity), it is more questionable using aggregate stock market data instead of data on single shares, since companies are likely to pursue very different dividend policies. But as Marsh and Merton have shown, *"it is (...) possible for aggregate dividends to exhibit stable and consistent time-series properties even if no such stability were found for individual firms."* Since, for theoretical and empirical reasons, the opposite is much less likely, it is advisable to use aggregate data if the empirical testing methodology strongly depends on capturing any systematic and stable element of dividend (policies) behaviour.[16]

2.1 The informational content of the dividend yield

Dividend yields, stock returns and dividend growth

We will begin with regressions that should reveal the information contained in the dividend yield for future stock returns and dividend growth. Equation (1.4) shows that the current

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[12] See Campbell (1987), p. 373.

[13] See Cuthbertson (1996), p. 129.

[14] Under risk-neutrality, asset returns should behave like martingales or random walks, respectively, which are unforecastable by definition. The neglect of time-variation in rational risk premia in a risk-averse world thus led to the long-held view that return predictability is synonymous to market inefficiency. See Kaul (1996), pp. 270-2.

[15] A more detailed description of the data is provided in the Appendix.

[16] See Marsh and Merton (1987), pp. 4 f.

dividend yield should predict future returns if the discount rates used by forward-looking investors actually depend on expected holding period returns for subsequent periods, and if these expectations do not deviate systematically, and too much, from realised returns. Since stock prices also depend on expected dividends, the dividend yield can only provide noisy measures of variation in expected returns, though, as Keim and Stambough put it, " *(...) whether this low signal-to-noise ratio destroys any ability of prices to predict returns is an empirical question.*"[17] The regressions for dividend growth are subject to the same omitted-variables problem because, in that case, expected stock returns introduce noise. To circumvent this problem, we also use the difference between returns and dividend growth as a single dependent variable.

Table 1

Long-horizon regressions of stock return measures on the log dividend yield

Regression equation: $\frac{1}{K}(x_{t+1}+...+x_{t+K}) = \alpha(K) + \beta(K)(d_t - p_t) + \varepsilon_{t+K,K}$

Estimation period with monthly data: December 1977 to June 1997

	Forecast horizon (*K*)					
	1	**3**	**12**	**24**	**36**	**48**
			$x_t = h_t$			
$R^2(K)$	0.001	0.013	0.052	0.102	0.120	0.352
$\beta(K)$	8.461	13.317	18.566	17.6113	16.240	17.389
t-value Newey and West	0.559	0.982	1.498	1.629	2.243	2.715
			$x_t = \Delta d_t$			
$R^2(K)$	0.046	0.108	0.229	0.166	0.143	0.107
$\beta(K)$	-18.770	-20.616	-15.058	-9.863	-7.217	-5.168
t-value Newey and West	-3.360	-3.327	-3.592	-2.761	-2.263	-1.659
			$x_t = h_t - \Delta d_t$			
$R^2(K)$	0.013	0.048	0.187	0.318	0.462	0.629
$\beta(K)$	27.231	33.933	33.624	27.474	23.457	22.557
t-value Newey and West	1.722	2.321	2.551	2.641	3.677	4.816

Notes: h is the annualised one-month continuously compounded stock return in per cent. Δd is the annualised one-month dividend growth rate in per cent. $(d - p)$ is the log dividend yield. $\alpha(K)$ (not shown) and $\beta(K)$ are the coefficients for the regression constant and the dividend yield, respectively, estimated by OLS. $\varepsilon_{t+K,K}$ are the error terms which are autocorrelated owing to data overlap for $K > 1$ under the null hypothesis of no predictability. Standard errors and t-values are corrected for serial correlation and heteroskedasticity in the equation error using the method of Newey and West (1987). Number of observations: 235 - (K-1).

Table 1 shows the regression results for each of the three dependent variables measured over a holding period (K months), ranging from one month to four years.[18] The regressions use monthly data, which means that data-overlap for the forecast horizons exceeding one month, induces

[17] See Keim and Stambaugh (1986), pp. 360 f.

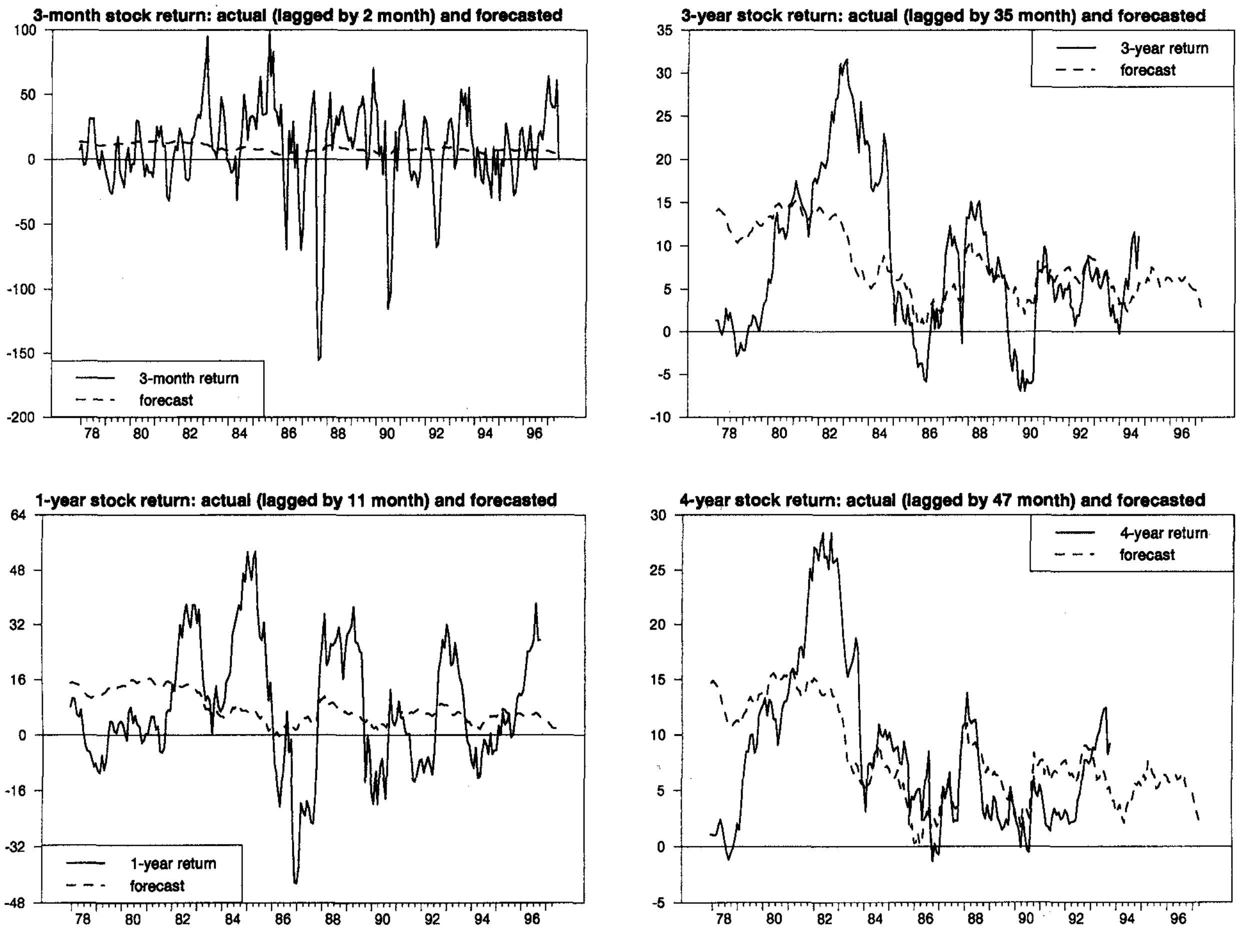
[18] The forecast horizons are chosen rather arbitrarily and follow the influential work of Fama and French (1988, 1989).

serial correlation of the error terms even under the null hypothesis of no return predictability (zero coefficient on the dividend yield). In this case, errors are correlated with K-1 previous error terms. But under alternative hypotheses, in which returns have a variable conditional mean, the serial correlation can in fact be arbitrary if dividend yields do not capture all of the variation in the conditional mean.[19] Additionally, since the regressor is only predetermined and not strictly exogenous, asymptotic distribution theory must be used to generate standard errors. The alternative t-statistics shown in the table for the null hypothesis of a zero coefficient are corrected for serial correlation and possible heteroskedasticity as suggested by Newey and West (1987) using a lag length of K-1.[20]

The upper part of Table 1 (see also Figure 3) summarises the main results for the stock returns regressions. The coefficient of determination (the $R^2(K)$ statistic) increases continuously with the forecast horizon, as do the t-values. The slope coefficients also increase from the one-month to the twelve-month horizon and remain roughly at that level for the longer forecast horizons. But statistical significance can only be attached to the 3-year and the 4-year return periods.[21]

Figure 3

Long-horizon regressions: stock returns and dividend yield



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[19] See Hodrick (1992), p. 360.

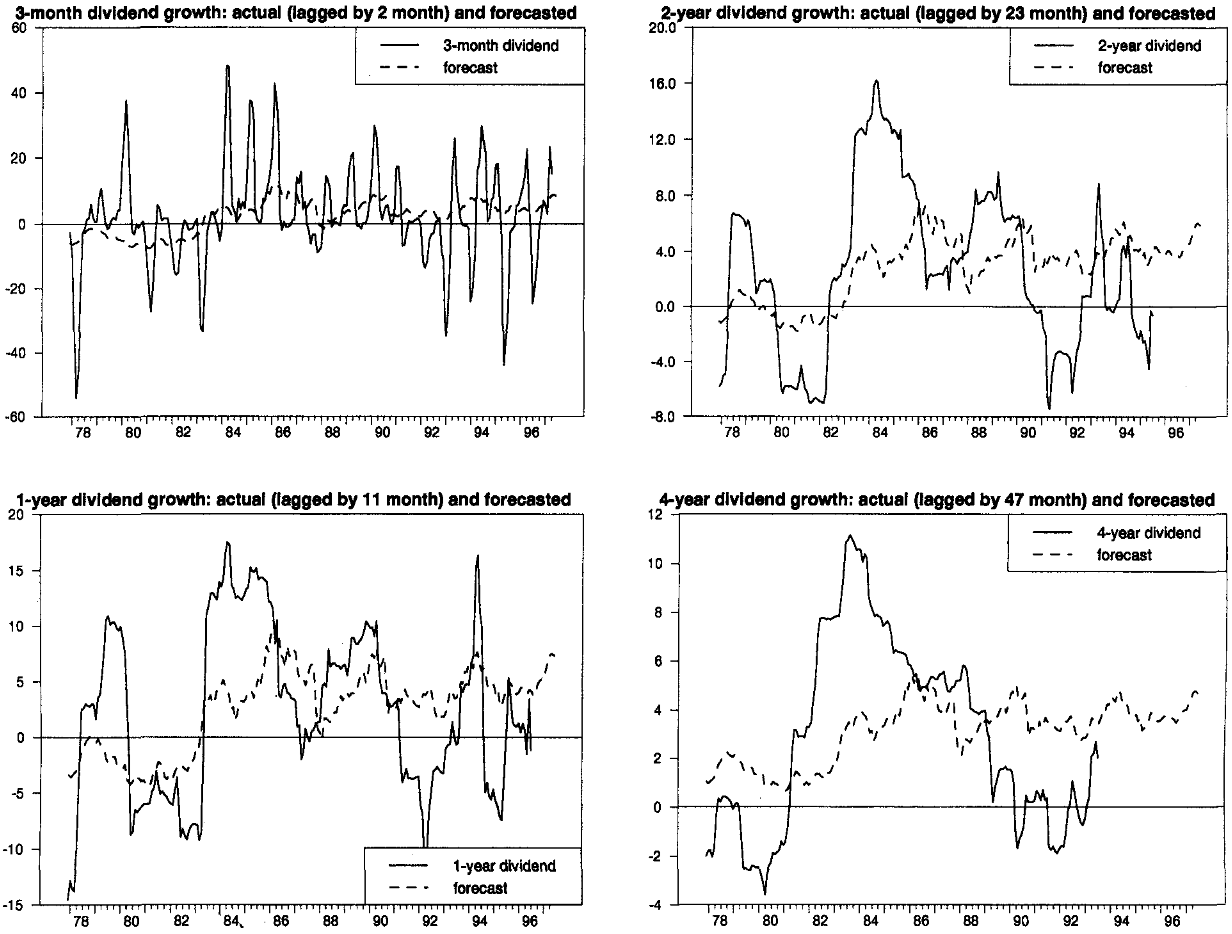
[20] These issues are discussed in more detail in Campbell et al. (1997), pp. 534-6.

[21] The results broadly conform to those for the US stock market, although our sample is much smaller, which weakens the comparability of results; see Fama and French (1988), p. 13 or Campbell et al. (1997), p. 269. The results for nominal returns are very similar to the results for real stock returns.

The R^2 statistics for the dividend growth regressions show, instead, a hump-shaped pattern and are much higher than in the stock returns case except for the 4-year horizon. They peak at the 1-year horizon with more than 20% of explanatory power (see Figure 4). What is more important is the high statistical significance of the slope parameters, particularly for the short to medium forecast horizons.

Figure 4

Long-horizon regressions: dividend growth [*ddiv*(*k*)] and dividend yield



The results for the combined returns variable (returns less dividend growth) are even more impressive. Although, by mere visual inspection, the time-series of this variable shows a very similar and volatile pattern as stock returns alone, the predictive power and the statistical significance of the slope coefficients are much higher for every forecast horizon (compare the results of the upper and the lower parts of Table 1, and see also Figure 5). The R^2 statistic increases to a remarkable 46 and 63% for the 3 and 4-year horizon, respectively. This comparison indicates that the noise introduced by dividend growth to the stock returns regressions is not negligible.

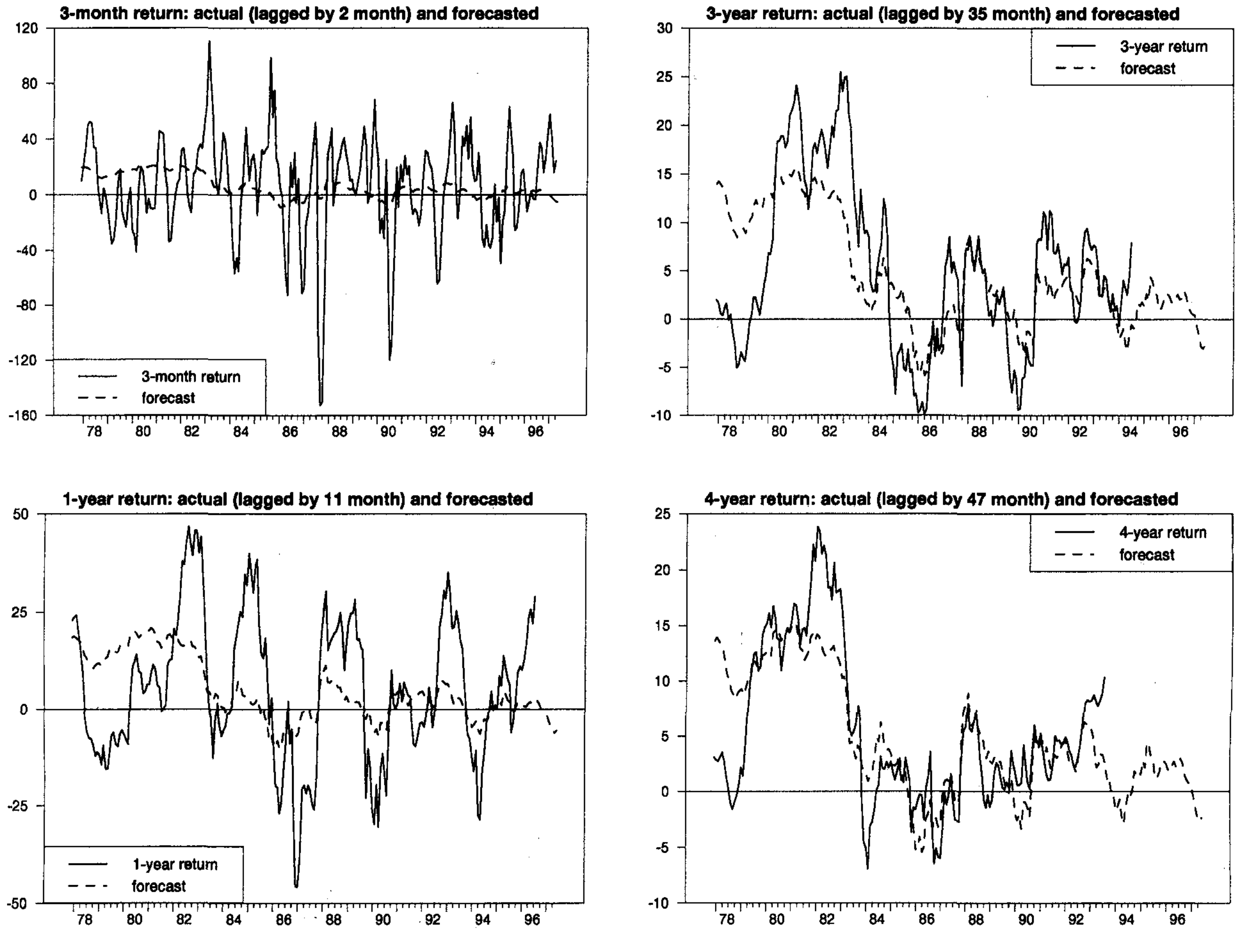
Although there are some serious doubts about the statistical reliability of long-horizon regressions, we interpret the results as providing sufficient preliminary evidence that future stock returns, and especially future dividend growth, contain predictable components which are reflected in the current dividend yield.[22] The fact that return predictability increases with the length of the holding period

[22] There are some general problems with long-horizon regressions in small samples. If the data are sampled more finely than the forecast interval, the error terms are autocorrelated at least of the order at which the data overlap. This autocorrelation is usually corrected for by use of some asymptotic distribution theory, in most cases with additional

considered may result from a better forecastability of the medium to long-term prospects of the economy (interest rates, business cycle positions).

Figure 5

Long-horizon regressions: combined returns $[r(k) - ddiv(k)]$ and dividend yield



From the point of view of market efficiency in terms of rational expectations (summarised by the RVF) the results of Table 1 indicate that dividend yields signal persistent time-variation in expected equilibrium returns, rejecting the long-held hypothesis that equilibrium returns are constant. The efficient markets hypothesis only postulates that abnormal returns are unpredictable,

correction for heteroskedasticity. But if the time span of data overlap is not small relative to the sample size, this approach is also flawed because there are not enough data points to reliably estimate the variance-covariance matrix. Monte Carlo simulations indicate that asymptotic standard errors can be very misleading in small samples (see Hodrick (1992), and Gerlach (1997), p. 164). An alternative is to use empirical standard errors using a bootstrapping procedure. But this method can only deal with biased standard errors. Another finite sample problem that puts into question the statistical reliability of long-horizon regressions derives from the fact that the independent variable, although predetermined with respect to the dependent variable, is stochastic and most likely correlated with past regression disturbances. This phenomenon leads to a finite-sample bias in the regression coefficients and the standard errors, *"(...) and the bias can be non-trivial even in samples of several hundred observations if the independent variable has both high autocorrelation and a high correlation with the past regression disturbance"* (Keim and Stambaugh (1986), p. 370). The dividend yield and term structure spreads – the independent variables used in this study – share at least the first property of being rather highly autocorrelated (i.e., highly persistent). But there are also some more theoretical problems caused by the very strong restrictions which rather simplistic models of the equilibrium returns process impose on the data. Modern theory suggests that the behavior of asset prices has much to do with the fundamental forces driving risk premia on the different assets. The assumption of constant risk premia provides a suitable starting point, but if risk premia actually play a significant role in asset pricing the econometrician most probably faces a serious omitted variables problem which biases coefficient estimates.

not that actual returns are unpredictable. High stock price volatility, as usually observed, is therefore compatible with persistent movements in rationally expected returns and need not indicate irrational investor behaviour.[23] But since return predictability could also result from irrational bubbles in stock prices, the question of whether the forecastability of stock returns is driven by rational economic behaviour or by animal spirits is still unresolved.[24] Further efficiency tests cannot fundamentally change this general conclusion but can only add evidence on the empirical plausibility of the rational valuation approach.[25]

Dividend yields and inflation

If we accept the view that stock prices are driven by expected equilibrium returns, it seems reasonable to ask whether the required rate of return includes a premium that compensates for inflation as expected over the holding period.[26]In that case, one could argue that the dividend yield should also have predictive power for future inflation.[27] But it has to be recognised that any empirical relationship between the two variables does not necessarily arise owing to an inflation premium in the dividend yield itself. If expected nominal dividend growth adjusts to inflation expectations in exactly the same way as the nominal discount rate does, the two effects on the dividend yield cancel out. The dividend yield can then be regarded as a real measure of stock returns and should not have any predictive power for future inflation unless expected real returns (including various risk premia) vary systematically with inflation expectations.

However, the regression results show for all forecast horizons high and significant slope coefficients which decrease with the horizon (see Table 2). The R^2 statistic is also always high, ranging from a minimum of 15% for the one-month period to a maximum of 54% for the 1-year horizon. The hump-shaped pattern of the R^2 statistic indicates that the forecast performance is best in a medium-term perspective (see also Figure 6).

How can this finding be interpreted in the light of the real nature of the dividend yield as explained above? We provide the following ad hoc explanation: First, assume dividend growth adjusts sluggishly to changes in the inflation environment. The expected dividend growth then falls short of the change in expected inflation. Second, if investors furthermore expect the central bank to raise (lower) short-term interest rates above (below) the upward (downward) shifts in expected or forecasted inflation, market participants will correspondingly require holding period returns which

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23 See Cuthbertson (1996), p. 129 and Campbell et al. (1997), p. 254.

24 In the case of bubbles, *"(...) dividend yields and expected returns are high when prices are temporarily irrationally low (and vice versa)"* (Fama and French (1989), p. 26).

25 To improve our understanding of the regression results in light of the rational valuation model we provide an illustrative example. When the log dividend yield decreases by 0.05 units from its long-term average (2.35% in logs) – which means a fall in the dividend yield of about twelve basis points – the average stock return tends to decrease by roughly 90 basis points over the next 4 years. This may be interpreted as follows: if investors require and expect a 90 basis points lower return on stocks, the log dividend yield will fall by 0.05 units. This in turn equals a 5% increase in the current stock price if dividends remain constant. The 25% increase from December 1996 until June 1997 (as measured with the price index used in this study) went along with a fall in the dividend yield of about 34 basis points. As predicted with the regression equation for 4-year returns, this fall is tantamount to a decrease in expected 4-year returns from 4.6 to 2.2%. This is a very low figure compared with average annualized stock returns of 8.5% over the past 18 years or so, but also relative to the level of short-term interest rates. Hence, if the forecast equation is not too biased, either rational investors are currently very risk prone regarding stock market investments, or economic agents behave irrationally, believing that the capital gains accrued over the recent months will continue or will at least not be reversed.

26 This does not preclude time-variation in real returns, which can be analysed separately, but is not the question of interest here.

27 The Fisher-effect can be analysed separately by running regressions between nominal stock returns and inflation or various proxies for inflation expectations. For some cross-country evidence see Solnik (1983).

increase (decrease) in excess of the inflation premium change. The net effect of the two offsetting channels through which changes in expected inflation influence share prices is to raise (lower) the current dividend yield, thus inducing a positive correlation between the dividend yield and future inflation.

Table 2

Long-horizon regressions of inflation on the log dividend yield

Regression equation: $\frac{1}{K}(\pi_{t+1}+\ldots+\pi_{t+K}) = \alpha(K)+\beta(K)(d_t - p_t)+\varepsilon_{t+K,K}$

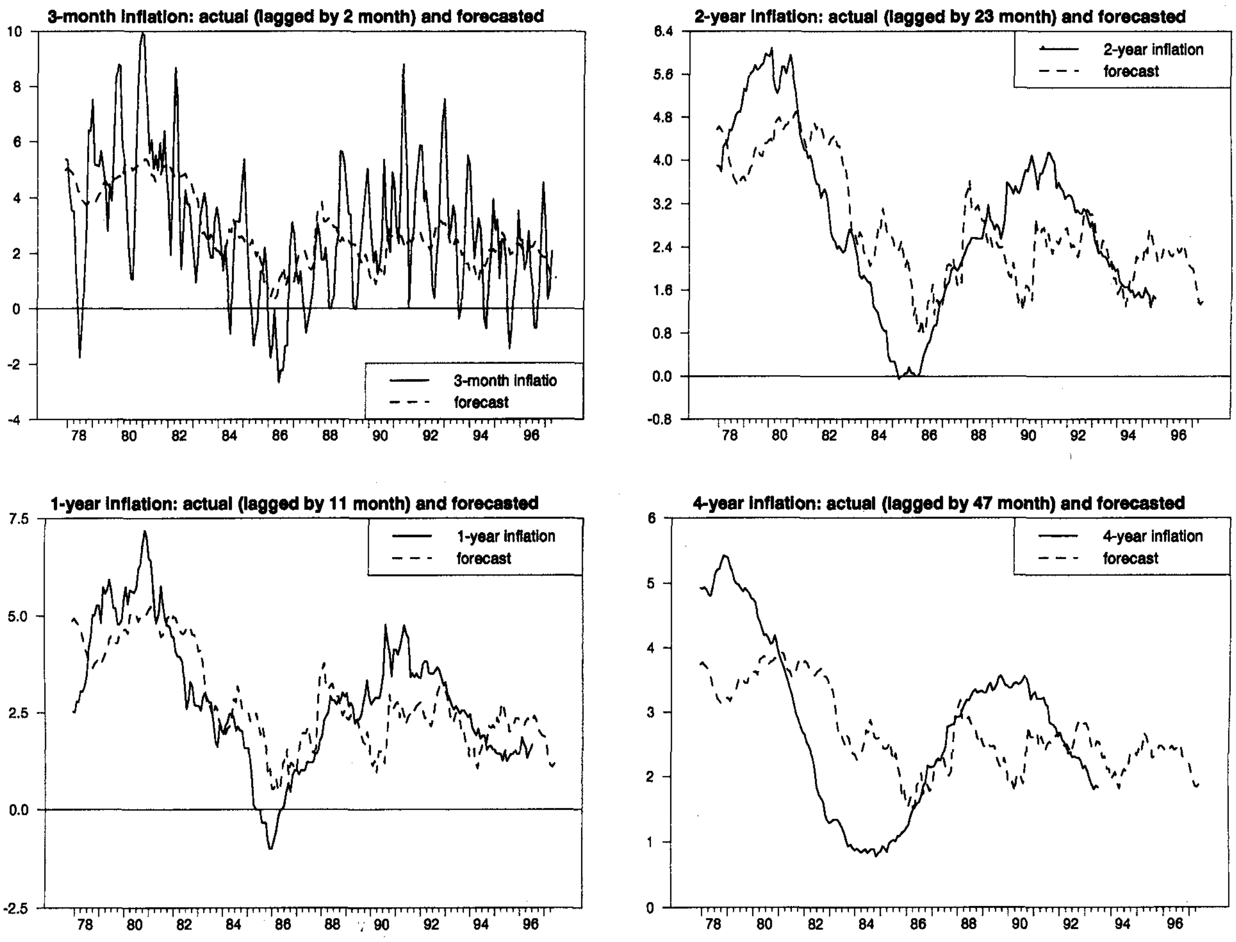
Estimation period with monthly data: December 1977 to June 1997

	Forecast horizon (*K*)					
	1	**3**	**12**	**24**	**36**	**48**
$R^2(K)$	0.151	0.277	0.542	0.470	0.367	0.241
$\beta(K)$	5.471	5.459	5.207	4.476	3.600	2.622
t-value Newey and West	6.436	5.857	5.342	3.847	3.117	2.366

Notes: π is the one-month continuously compounded rate of consumer price inflation. $(d - p)$ is the log dividend yield. $\alpha(K)$ (not shown) and $\beta(K)$ are the coefficients for the regression constant and the dividend yield, respectively, estimated by OLS. $E_{t+K,K}$ are the error terms which are autocorrelated owing to data overlap for $K > 1$ under the null hypothesis of no predictability. Standard errors and t-values are corrected for serial correlation and heteroskedasticity in the equation error using the method of Newey and West (1987). Number of observations: 235 - (K-1).

Figure 6

Long-horizon regressions: inflation and dividend yield



But some words of caution have to be added. Inflation and the dividend yield are highly persistent variables. According to standard unit-root tests, both variables can be regarded only as borderline stationary or near-integrated. From a mere statistical point of view, it is thus possible that the high R^2 statistics result from stochastic trends in the data and are thus spurious.

2.2 The information content of the term structure spread

The term structure spread and short-term interest rate changes

According to the expectation hypothesis with rational expectations (*EH-RE*), the spread is an optimal predictor for future changes in short-term interest rates. The spread should equal a weighted average of expected short-rate changes over the life of the long bond plus a constant risk premium. Referring to the long-horizon regression methodology, one can test the forecast accuracy by constructing the perfect foresight spread, $S_t^{n(pf)}$, for each bond maturity n from ex post values of short-rate changes as:

$$S_t^{n(pf)} = \sum_{i=1}^{n-1}(1 - i/n)E_t \Delta r_{t+i} + \phi^{(n)} \tag{2.1}$$

Table 3

Long-horizon regressions of the perfect foresight spread on the actual spread

Regression equation: $S_t^{n(pf)} = \alpha(n) + \beta(n)S_t^n + \varepsilon_t^n$

Estimation period with monthly data: September 1972 to June 1997

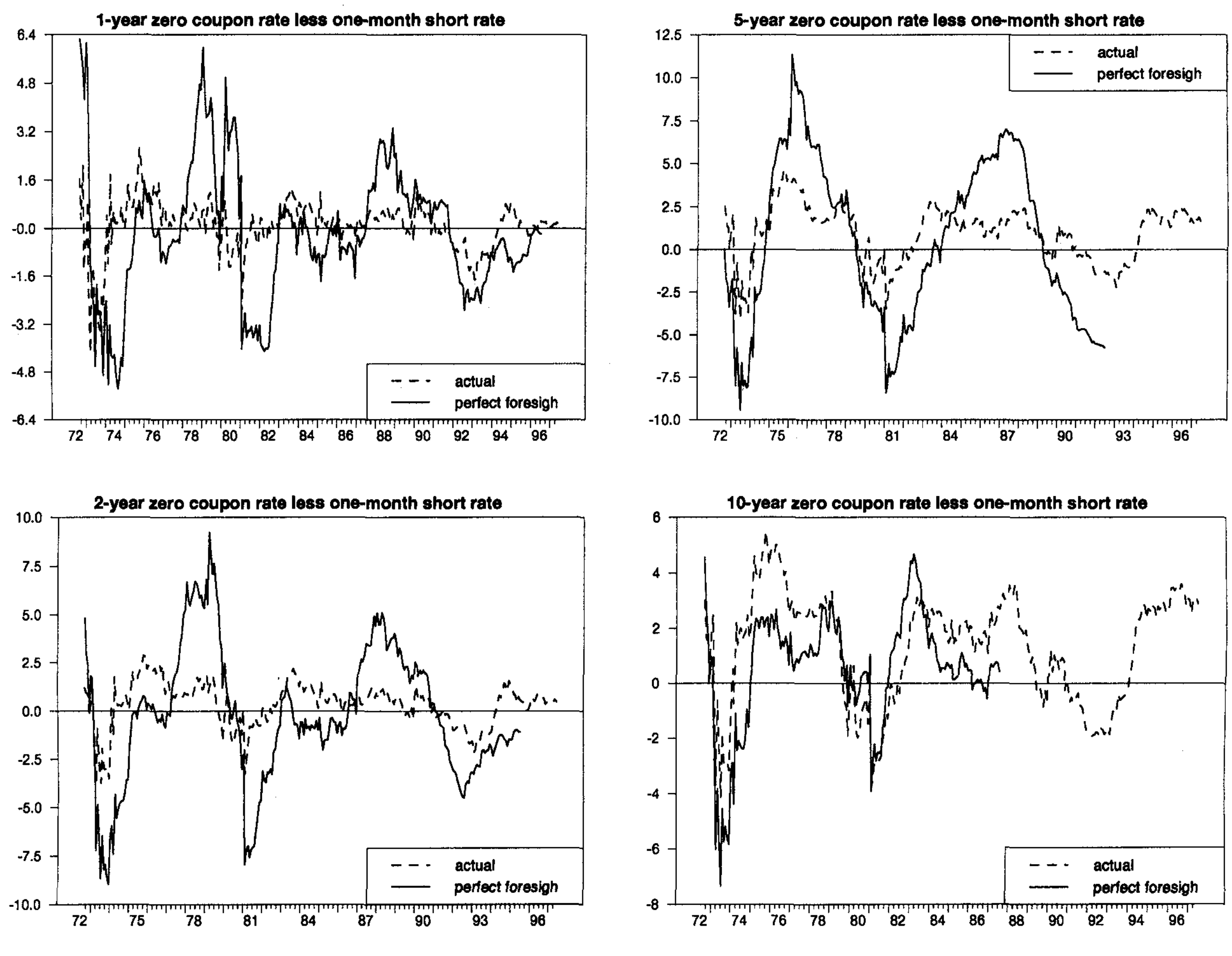
	Long-bond maturity in years (*n*/12)									
	1	**2**	**3**	**4**	**5**	**6**	**7**	**8**	**9**	**10**
Term premium $\phi(n)$	0.07	0.34	0.58	0.78	0.93	1.05	1.14	1.21	1.26	1.31
R^2	0.13	0.33	0.40	0.51	0.70	0.71	0.59	0.43	0.47	0.50
$\alpha(n)$	-0.21 (0.36)	-0.76 (0.64)	-1.02 (0.77)	-1.30 (0.72)	-1.75 (0.53)	-1.56 (0.62)	-0.76 (0.88)	-0.13 (0.79)	-0.34 (0.27)	-0.56 (0.65)
$\beta(n)$	0.89 (0.19)	1.69 (0.38)	1.86 (0.37)	2.08 (0.23)	2.36 (0.17)	2.11 (0.22)	1.51 (0.39)	0.94 (0.40)	0.84 (0.08)	0.73 (0.14)
H0: $\beta(n) = 1$	0.54	0.07	0.02	0.00	0.00	0.00	0.19	0.88	0.05	0.06
H0: $\alpha(n) = 0$, $\beta(n) = 1$	0.72	0.03	0.00	0.00	0.00	0.00	0.40	0.85	0.14	0.00
Variance ratio (*VR*)	0.40	0.34	0.34	0.34	0.35	0.40	0.51	0.71	0.80	0.93

Notes: $S_t^{n(pf)}$ is the perfect foresight spread as defined in equation (2.1) using the respective term premium as given in the first line of the table. S_t^n is the actual spread between the n-period (in months) bond and the one-month interest rate. $\alpha(n)$ and $\beta(n)$ are the coefficients (standard errors in brackets) for the constant term and the actual spread, estimated by OLS. ε_t^n are the error terms which are autocorrelated of order n-1 due to data overlap. Standard errors are corrected for serial correlation and heteroskedasticity in the equation error using the Newey and West (1987) method. The values shown for the hypothesis tests are p-values; the test statistic for the Wald-test is distributed as $\chi^2(df)$ with $df = 1$ and 2 degrees of freedom. The variance ratio is defined as the sample standard deviation of the actual spread, divided by the standard deviation of the perfect foresight spread. Number of observations: 298 - (n-1).

and regress it on the actual spread and a constant. We do this for spreads between long-bond zero-coupon rates and the one-month interest rate on the interbank money market in Frankfurt. The long-bond maturities tested range from 1 year (n = 12 months) to 10 years (n = 120 months). In constructing the perfect foresight spread we face the problem of how to get an estimate of the term premium. We use a common but rather crude method and estimate the term premium for each maturity by the difference in the sample means of the respective long rate and the short-term interest rate.[28] As can be seen from the first line of Table 3, the estimated term premia increase with bond maturity. This is not compatible with the conventional interpretation of the *EH* which assumes *constant and equal* term premia for all maturities. Instead, the relevant hypothesis to be tested is the liquidity preference hypothesis, which exactly adds to the *EH* the assumption of term premia increasing with bond maturity. For the sake of simplicity, we subsume the liquidity preference hypothesis under the notion *EH*.

Figure 7

Long-horizon regressions: perfect foresight spread and actual spread



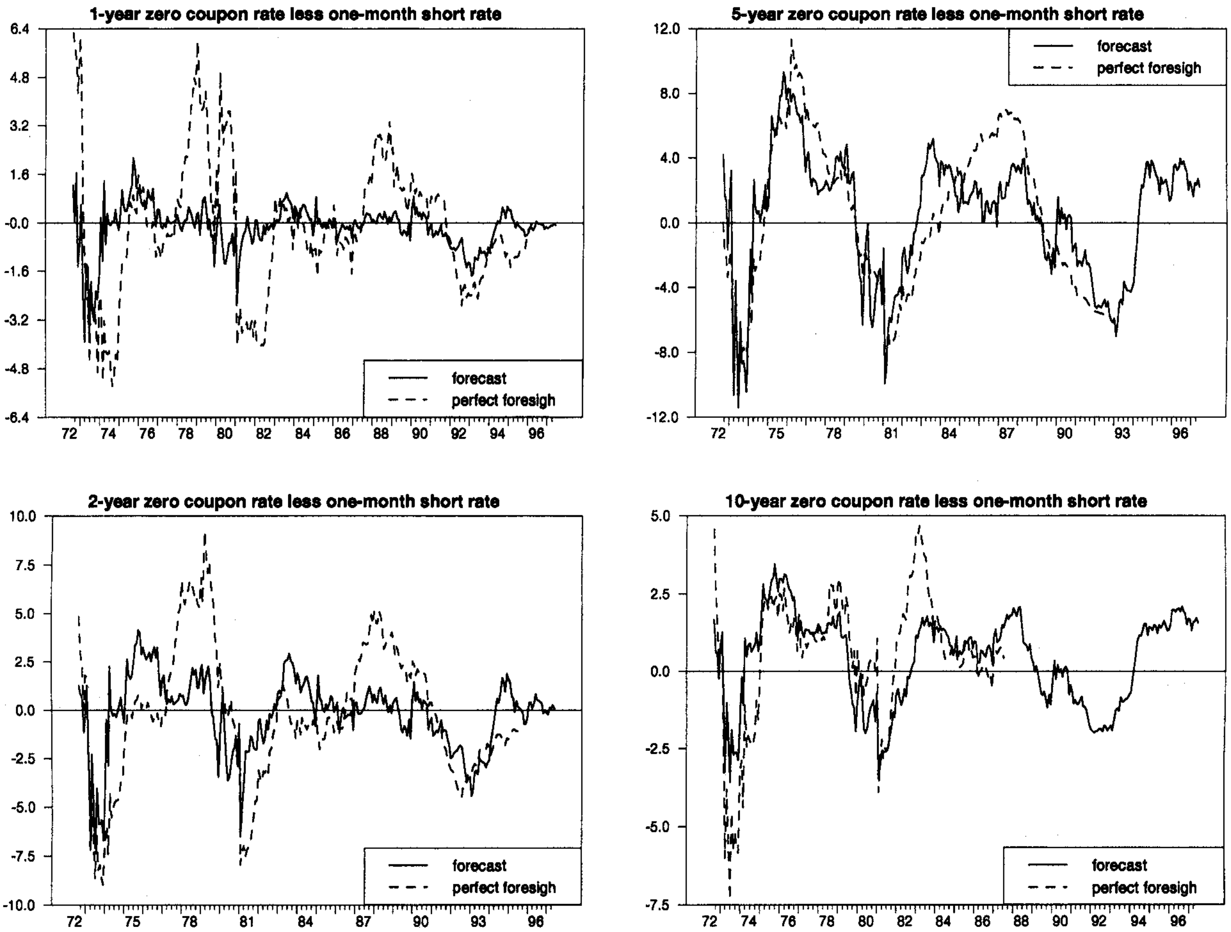
The R^2 statistic is rather high for all maturities but the one-year horizon. It peaks at the medium-term maturities of 5 and 6 years, at about 70%. The slope coefficients show a more pronounced hump-shaped pattern with the highest value of 2.36 for the 5-year maturity. Thus, high (low) R^2 statistics tend to be associated with high (low) slope coefficients. Taken together, this suggests that investors can reliably predict only medium-term, but not very near-term, developments of future short rates, which may be based on better medium-term forecastability of real activity and

28 See, for example, Shiller (1989), p. 225.

inflation. But although the direction of change has been anticipated quite well, the magnitudes of the medium term interest rate changes have been underestimated, causing the slope coefficients to get significantly larger than one to improve the regression fit (compare Figures 7 and 8). This explanation may hold at least within the given sample, which includes highly volatile periods, such as the oil-price shocks and German reunification. However, the interest rate shocks associated with these exceptional phases were only temporary and vanished or cancelled out after the medium term. Thus, for times to maturity of 8 to 10 years the accumulated short-rate changes are much lower, the slope coefficients are around one, and the relative standard deviations of the actual and the perfect foresight spreads (the variance ratio) approach unity. But it has to be conceded that differences in the slope coefficients may also arise from the influence of omitted variables, especially those factors which may introduce time-variation in the term premia.

Figure 8

Long-horizon regressions: perfect foresight spread and its forecast



This regression framework also forms a basis for testing market efficiency or the *EH* using a rather strong definition of rational expectations (*EH-RE*). It assumes that investors can forecast future short-rate changes perfectly save a pure white noise error which is orthogonal to all information at time t (the forecast origin):

$$\Delta r_{t+i} = E_t \Delta r_{t+i} + \eta_{t+i} \tag{2.2}$$

with i = 1, ..., n-1. Substitution into (2.1) leads to the testable hypothesis that the perfect foresight spread should equal the actual spread (its optimal predictor); differences between the two should be purely random and uncorrelated with all information available at time t or earlier (to which the actual spread itself belongs, too):

$$S_t^{n(pf)} = S_t^n + \varepsilon_t^n \tag{2.3}$$

The regression equation in Table 3 represents the appropriate testing framework. Under hypothesis (2.2) the regression error is a moving average process of order (n-1) for monthly data:

$$\varepsilon_t^n = \sum_{i=1}^{n-1} (1 - i/n)\eta_{t+i} \tag{2.4}$$

The expected value of the compound forecast error is still zero, but successive errors are autocorrelated and possibly heteroskedastic. The standard errors for the regression coefficients are therefore again corrected for serial correlation and heteroskedasticity, using the Newey-West method.[29] The *EH-RE* or efficient market hypothesis implies the restrictions $\beta(n) = 1$ and $\alpha(n) = 0$. Table 3 shows the p-values for Wald-tests of the first restriction (fourth line) and of both restrictions together (fifth line). The best results from the efficient market view are for the 1-year and the 7-year to 9-year maturities with sufficiently high p-values for both restriction sets. Particularly for the medium-term maturities, the spread is a biased (with slope coefficients much above one, the value implied by the efficient markets hypothesis), although a better predictor of future short-rate changes.

This model-consistent performance of the longer maturities also shows up in the variance ratios, which are much higher than for the shorter maturities and approach unity for the 10-year maturity. As can be derived from equation (2.3) and the null hypothesis of *RE*, the variance (or the standard deviation) of the perfect foresight spread must always be higher than the variance of the actual spread.[30] This is actually the case for all maturities, but since the variance ratio (actual to perfect foresight spread) approaches one with decreasing variance of forecast errors, a high (low) variance ratio indicates low (high) forecast error variances. Hence, the accumulated long-run forecast errors tend to be significantly lower than errors summed over shorter time periods. This in turn confirms our conjecture, above, that the cancelling-out of temporary strong interest-rate movements over the longer periods reduces the bias in the slope coefficient and hence weakens evidence against the efficient market hypothesis.

However, there are still some more fundamental doubts about the appropriateness of using perfect foresight measures of expectations as the basis for testing market efficiency. This very strong hypothesis of *RE* assumes that agents can forecast with 100% accuracy, regardless of any unforeseeable special events that occur during the sample. An alternative, ex ante oriented, approach tries to find a suitable (multivariate) time-series representation of the data and expectations generating process and to draw inferences about market efficiency from forecasts based on such models.[31]

Term structure spread and inflation changes

The Fisher theorem states that the current nominal interest rate of a bond in equilibrium equals the expected real interest rate plus the (annualised) expected rate of inflation over the life of the bond. The real rate also contains any risk premium required by investors. If this relation holds and

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29 See Cuthbertson (1996), p. 325.

30 See Cuthbertson (1996), p. 138.

31 The so-called Campbell and Shiller (1987) approach provides some metrics to test market efficiency in this context. For some exemplary evidence on the German bond market see Gerlach (1996). Domanski and Kremer (1997) apply this approach to the German stock market.

if the real interest rate is constant, then the spread between the interest rates of an m-year and a j-year bond should exactly correspond to the (annualised) difference in expected inflation m years and j years ahead, respectively. Hence it makes sense to use term structure spreads as indicators of changes in inflation expectations held by market participants. In a recent study, Schich (1996) analyses the predictive content of spreads regarding future inflation changes by using zero-coupon rates for the German government bond market. We refer to this study for the details and show slightly updated results for the long-horizon regressions in Table 4 (see also Figure 9).[32]

Table 4

Long-horizon regressions of inflation changes on spreads

Regression equation: $\Delta\pi_t^{(m,1)} = \alpha(m,1) + \beta(m,1)S_{t-1}^{(m,1)} + \varepsilon_t^{(m,1)}$

Estimation period with monthly data: September 1972 to June 1997

	Longer-bond maturity in years (m)								
	2	**3**	**4**	**5**	**6**	**7**	**8**	**9**	**10**
R^2	0.040	0.113	0.203	0.305	0.362	0.357	0.348	0.330	0.273
$\alpha(m,1)$	-0.188	-0.441	-0.724	-1.022	-1.267	-1.322	-1.468	-1.540	-1.528
	(0.120)	(0.210)	(0.282)	(0.349)	(0.389)	(0.466)	(0.400)	(0.373)	(0.388)
$\beta(m,1)$	0.268	0.450	0.633	0.801	0.893	0.851	0.816	0.746	0.646
	(0.164)	(0.180)	(0.135)	(0.120)	(0.125)	(0.143)	(0.125)	(0.116)	(0.132)

Notes: $\Delta\pi_t^{(m,1)} = \pi_t^{(m)} - \pi_t^{(1)}$ is the difference between the annualised one-year and m-year-ahead rate of inflation defined as $\pi_t^{(j)} = 100/j(p_{t+12j} - p_t)$ with $j = 1, \dots m$ and p the log of the German consumer price index, for monthly data. The spread variable is simply defined as the difference between the zero-coupon rate for an m-year and the one-year bond, i.e., as $S_t^{(m,1)} = z_t^m - z_t^1$. $\alpha(m,1)$ and $\beta(m,1)$ are the coefficients (standard errors in brackets) for the constant term and the spread variable, estimated by OLS. $\varepsilon_t^{(m,1)}$ are the error terms which are autocorrelated of order ($12m$ - 1) due to data overlap. Standard errors are corrected for serial correlation and heteroskedasticity in the equation error using the Newey and West (1987) method. Number of observations: 298 - ($12m$ + 1).

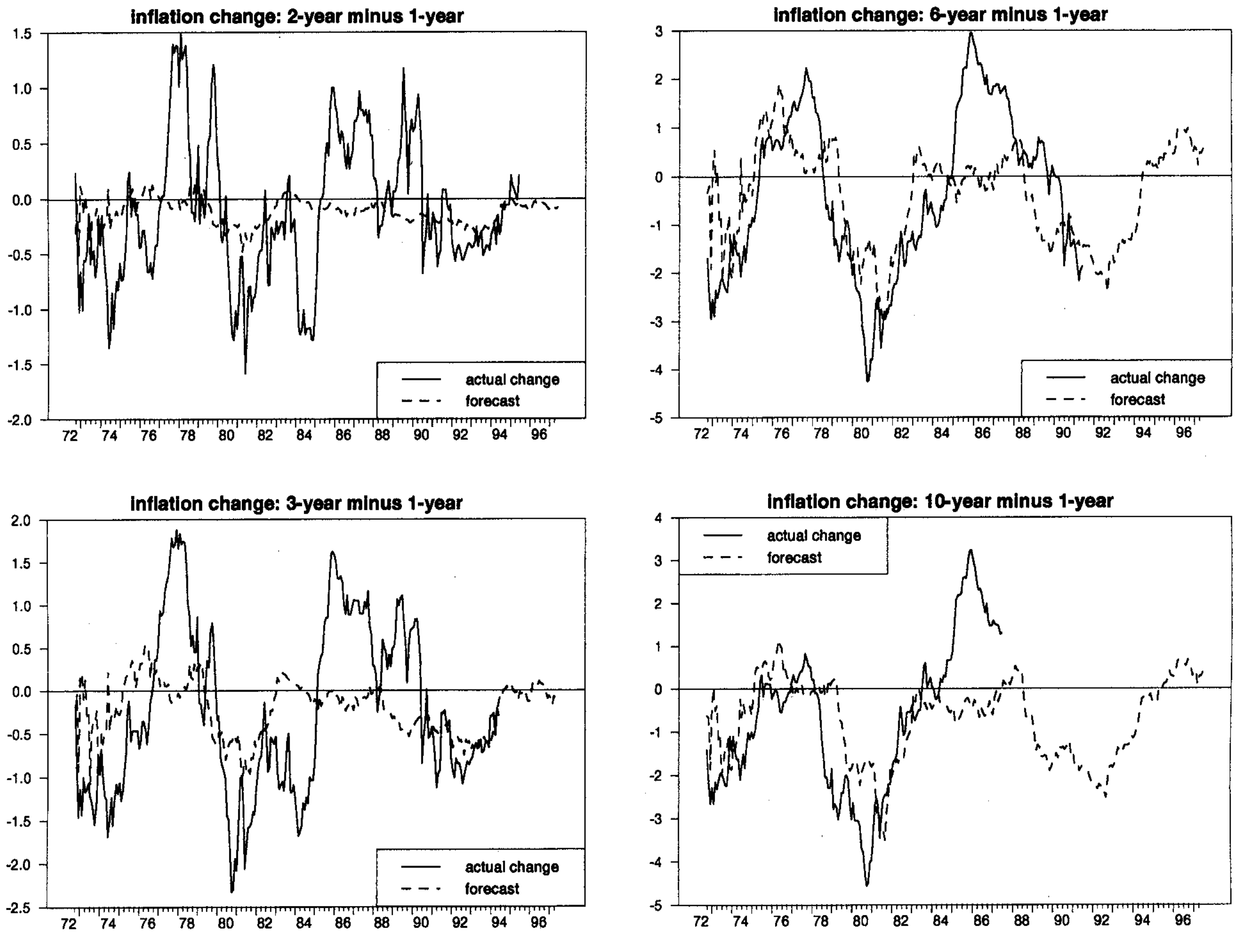
The empirical evidence suggests that especially the medium to long-term segments of the German term structure of interest rates have significant forecast power for future inflation changes over the respective horizons with R^2 statistics of up to about 40%. This, in turn, favours the interpretation that market agents can anticipate longer-term developments of inflation better than short-term changes. The values of the slope coefficients show the hump-shaped pattern sometimes observed in the other long-horizon regressions above. For the longer maturities, they are in the neighbourhood of one, the value which would obtain if the strong *RE* assumption were imposed on inflation expectations. These general results are robust to the use of either zero-coupon rates or yields to maturity.[33]

32 See Schich (1996), pp. 39-50. See also Gerlach (1997) for a brief survey of related studies and the origins of the testing methodology.

33 See again Schich (1996), who compares the results obtained for both interest rate measures.

Figure 9

Long-horizon regressions: inflation and spread forecast



3. Implications for monetary policy

3.1 Impact of monetary policy on asset prices

According to the rational valuation approach, monetary policy can influence asset prices by two channels. First, the price of long-term assets like stocks and bonds reflects agents' expectations about the course of the short-term interest rate which may serve as a benchmark for equilibrium one-period returns. These short-term returns are used for discounting the assets' future streams of income. As monopolistic supplier of base money, the central bank can – at the operational level – determine short-term interest rates and thus affect asset prices via agents' expectations about the future path of money market rates. Second, since in general the nominal long-term returns which investors require to hold an asset in their portfolios should contain an inflation premium, it is the long-horizon perspective about future inflation that influences today's prices of long-term assets. At the strategic level, however, monetary policy controls inflation in the longer run. Hence, monetary policy has a strong impact on asset prices by affecting agents' inflation expectations over longer horizons.

But both channels are merely two sides of the same coin since in equilibrium successive short-run returns simply have to add up to long-run expected returns. The long and short-term perspectives are interlinked by the central bank's reaction function as perceived by economic agents.

A change in inflation expectations, for example, should cause a shift in the path of expected short-term rates and vice versa. This link has clear implications for the way monetary policy decisions affect the level of and changes in asset prices. By reducing short-term rates below equilibrium level, the central bank may increase stock prices and the term spread if long-term expectations remain unaffected. But this only occurs if the central bank measure conforms to the monetary policy regime implied by the reaction function which agents use to determine their expectations about future inflation and short-term interest rates. If the measure comes as a surprise and does not fit to previous experiences with monetary policy, there always exists the danger that asset prices react in a way which counteracts the intentions of central bankers. Short-term fluctuations of asset prices – their volatility – in this view depend on how often and to what extent expectations have to be revised by market participants.

A first conclusion from this is that a predictable monetary policy makes it easier for economic agents to form expectations. First, through an unambiguous obligation to the goal of price stability monetary policy provides a nominal anchor for inflation expectations over longer horizons. Second, a transparent strategy establishes a link between this strategic level and the operational level reflected in short-term interest rates. Under this conditions, it is reasonable for market participants to assume that short-term (policy-determined) rates might fluctuate significantly in the short run (in order to contain inflationary pressures and to make real "monetary" rates conformable to expected changes in real "capital" rates), but should return to a "normal" level in the medium run. Third, monetary policy should be able to smooth market volatility by reducing uncertainty of future rate changes. This, again, is a facette of a transparent strategy, but is also related to the implementation of monetary policy. If money market rates fluctuate by chance or in an undesired manner owing to unexpected changes in banks' liquidity, there can be volatility spill-overs to other financial markets.[34]

From this point of view, the empirical results presented above can be seen as an indication that monetary policy in Germany has been able to provide a relatively reliable medium-term orientation, thus facilitating the process of expectation formation regarding inflation and short-term interest rates. The fact that the forecast performance of the dividend yield with respect to future inflation is better at shorter horizons than in the case of the term structure spread may indicate that other factors which determine stock returns dominate the influence of inflation, especially over longer horizons. That is, the noise introduced by the omitted variables in the forecast equations for inflation (changes) is probably stronger for the dividend yield regressions. Furthermore, the results support the view that short-term expectations about stock returns and money market rates are often subject to disappointments reflecting unprecedented macroeconomic shocks. In the short run, these shocks can have a very strong and unexpected impact on inflation rates and the path of short-term interest rates which renders econometric analysis – using either ex post data or ex ante measures of the variables to be forecasted – more difficult.

3.2 The use of asset prices as monetary policy indicators

The empirical evidence presented in this paper shows that the dividend yield and the term structure spread contain useful information about future stock returns, dividend growth, short-term interest rate changes and inflation (changes) as expected by market participants, at least over medium-term horizons. At a first glance, this seems to support an outstanding role for financial market prices as indicators for monetary policy. However, although the regression fit is in most cases impressive according to standard metrics, the forecast errors are generally rather high from an operational point of view. Thus, policy makers face a lot of uncertainty if they try to evaluate whether any change in the indicator variable reflects shifts in agents' expectations or, instead, the influence of other factors omitted from the forecasting equation. Moreover, from a strategic perspective, it is crucial that monetary policy still relies on an "external" anchor and not on market expectations themselves.

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[34] See Schmid and Asche (1997).

The anchoring of expectations about monetary policy can probably best be achieved by a strong and credible commitment to long-term price stability. The respective long-term inflation goal is then given a heavy weight in any reaction function which economic agents use in forming their expectations about the future course of short-term interest rates.

By instead linking monetary policy decisions to market expectations, the form of expectations about inflation and, connected to that, the future path of short-term interest rates becomes self-fulfilling and could lead to policy instability and hence inflation instability.[35] This makes room for speculative attacks in financial markets and jeopardises the credibility of the central bank.

Independently of the danger of sliding into a vicious circle, putting more weight on market expectations could be interpreted as a shift in the monetary policy regime by market participants. This makes it difficult for the central bank to assess the stance of monetary policy because market indicators become less reliable (which should show up in coefficient changes in the forecasting equations) and other indicators (as, for example, the money stock) may lose their indicator properties owing to changes in the behaviour of market participants. Finally, the central bank could end up in a situation in which it is impossible, or at least rendered more difficult, to stabilise expectations just because monetary policy has been geared to market expectations. All this suggests, as Woodford convincingly argued, that modelling structural relationships, including the monetary policy reaction function, is unavoidable in order to make more reliable inferences about the indicator quality of a financial market variable and to assess its usefulness for monetary policy purposes.[36]

Appendix: Data description

The monthly stock price and dividend series used in this study are calculated by the Federal Statistical Office up to June 1995. The computations are based on a fictitious share having the face value of DM 100. The stock price series is the arithmetic mean of the end-of-month prices of all the shares of public limited companies officially listed on German stock exchanges (stock prices of each company are previously multiplied by a factor which raises or lowers its face value to DM 100). The series is thus equivalent to an equally-weighted stock price index. The dividend series is calculated correspondingly. However, the monthly dividend (excluding tax credit) of each share is the dividend as last paid out. The dividend yield (in per cent per annum) is defined as the ratio of dividends to stock prices multiplied by one hundred. While the stock price series is available for the period from January 1960 to June 1995, the dividend series only begins in November 1977. Both series are published in Deutsche Bundesbank, Capital Market Statistics, Statistical Supplement to the Monthly Report 2, Table IV.2. Complementary series for the period from July 1995 to the present are calculated by the Deutsche Börse AG. But as the number of stocks included in the calculation is reduced (only ordinary and preference shares officially listed on the Frankfurt stock exchange of companies domiciled in Germany are included) a statistical break occurs which is accounted for in the empirical analysis.

The interest rates representing the German term structure are estimated zero-coupon rates. They are estimated from the prices of listed coupon bonds issued by the Federal Government. For a detailed description of the estimation procedure see Deutsche Bundesbank (1997). The monthly series comprise end-of-month data as published in Deutsche Bundesbank, Capital Market Statistics, Statistical Supplement to the Monthly Report 2, Table II.7e).

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35 See Woodford (1994), p. 104.

36 See Woodford (1994), p. 112 f.

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The information content of financial variables for forecasting output and prices: results from Switzerland

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Introduction

Central banks need reliable forecasts of output and prices to conduct monetary policy. Forecasts of prices are important because central banks aim at delivering price stability in the long run. Forecasts of output are necessary because, under certain conditions, central banks may find it helpful to influence the business cycle and to stabilise output in the short run. Several financial variables have long been used as important information variables to forecast output and prices. Traditionally, a monetary aggregate, like M0, M1 or M2, has been the key variable for many central banks. Several authors recently documented the decline in the forecasting ability of money, especially in the case of the United States.[1] Thus, the relation between money and prices and between money and output has become loose. At the same time, many of these authors suggested that interest rates and interest rate spreads dominate money as information variables.[2] However, it is also possible that other financial variables and asset prices contain important information to forecast prices and output. This is especially of interest because of the large movements in financial variables, such as exchange rates and stock market prices, in recent years. Movements of asset prices and financial variables may reflect expectations of market participants. These expectations usually have a strong impact on changes in output and prices. Movements of asset prices and financial variables, however, may also be the consequence of large portfolio shifts and financial innovations. Such shifts in the financial structure of a country are important because they signal possible changes in the money demand function. Generally, large movements in financial variables may lower the information content of money and render a monetary policy based on monetary aggregates more difficult to pursue. Exploiting the information from other financial variables can alleviate this problem.

In 1992, the Swiss National Bank started to pursue a more flexible monetary policy by announcing a medium-term target for its monetary base. This allows the use of a broader spectrum of information variables. The monetary policy of the Swiss National Bank may not exclusively depend on the development of the monetary base, particularly in the short run. Nevertheless, the Swiss National Bank considers the monetary base as the most important information variable for prices in the long run and therefore formulates a medium target for base money. In the short run, alternative indicators become more important for policy decisions, independently of whether the Swiss National Bank follows a policy of monetary targeting or a policy of inflation targeting.[3]

The aim of this paper is therefore to evaluate in what respect financial variables other than monetary aggregates help to forecast output and prices. As pointed out by Sims (1972) and by Friedman and Kuttner (1992), a financial variable is a useful information variable for forecasting output and prices if fluctuations in this variable, not only predict fluctuations in prices and output in general, but also movements which are not foreseeable from past fluctuations in output and prices. As

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1 See, for example, the influential papers by Friedman and Kuttner (1992, 1993 and 1996) and Friedman (1996).

2 See also Bernanke (1990).

3 Some countries started to use a monetary conditions index as an information variable. Lengwiler (1997) shows that such an index does not deliver superior information compared to a monetary aggregate in Switzerland.

long as a variable "Granger'' causes money and prices, it can be used by the central bank as an information variable, independently of the exact nature of the causation. However, an information variable is most useful if its predictive power remains stable over a long period of time.

In the sections below, I use vectorautoregression methodology in order to analyse the information content of various variables and systems. Besides the monetary aggregate M2, the analysis is applied to a broad set of different financial variables,[4] including the bond interest rate, the spread between the short and the long-run interest rate, the trade-weighted nominal exchange rate index and the stock market index. The focus is thus to check what type of financial variable can potentially be important. Further research could, for example, look closer at a set of different exchange rates or at different interest rates.

The analysis shows that money and the exchange rate index are the most important information variables of those considered. Money (M2) is especially helpful in predicting output. The systems including money keep their forecasting superiority over time, although it has recently become more difficult to predict output. The exchange rate index has a predictive content about prices. However, this forecasting information erodes over time, so that, at the end of the considered samples, forecasts of prices based only on past output and prices outperform all other forecasts.

The paper is organised as follows: In Section 1, the in-sample predictive content of the set of financial variables is analysed by considering variance decompositions. Section 2 looks at the out-of-sample forecasting ability. In Section 3, the change of the predictive content is analysed and the last section concludes.

1. In-sample predictive content

This section analyses the in-sample information content of variables for predicting output and prices by estimating vectorautoregressions (VAR) of various systems.[5] I start by considering the integration order of the variables included in this study. The augmented Dickey-Fuller test and the Phillips-Perron test indicate that all variables are integrated of order I(1), with the exception of the spread between the long and the short-run interest rate. Although not completely clear-cut, the tests point toward stationarity of the spread.[6] VARs with integrated variables are usually estimated with differenced data. However, differencing leads to a loss of information if a cointegrating relation is present.[7] Instead, VARs in levels preserve possible cointegrating relationships among the variables without explicitly imposing a specific cointegration vector. Therefore, the Johansen procedure is applied for the basic systems considered below in order to test for cointegration between the variables. The results indicate that, indeed, the null hypothesis of no cointegration between the variables of the systems considered in the subsequent analysis can be rejected.[8]

Since there is possible cointegration among the variables, I estimate different vectorautoregressions with variables in levels. As pointed out by Sims, Watson and Stock (1990) and

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4 Note, however, that the set of financial variables available for analysis is rather limited in Switzerland.

5 A methodological alternative would be to use vector error correction models.

6 The results of the unit root tests are not reported here. For trending variables, the regressions of the test include a constant term and a time trend. For non-trending variables, i.e., interest rates and spreads, only a constant is included.

7 See Sims, Watson and Stock (1990) and Hamilton (1994).

8 The cointegration results are not reported here. They are in line with the findings from other studies based on similar data. See, for example, Flury and Spörndli (1994).

Hamilton (1994), standard Granger causality tests (F-tests on all lags of the same variable) are not valid if a VAR consisting of I(1) variables is estimated in levels. However, no statistical problems exist for the computation of variance decompositions. Furthermore, as put forward by Thoma and Gray (1994), F-statistics can be misleading indicators of causality, because the effect of one variable on another may be transmitted through a third variable. In addition, F-tests only refer to the one-quarter-ahead prediction while variance decompositions allow for predictions over a longer horizon. Variance decompositions are capable of measuring the quantity of the predictive content of a variable whereas F-tests only indicate whether a variable has any information content at all.[9]

The regressions used in this section of the paper take the form:

$$x_t = D(L)x_{t-1} + \varepsilon_t \tag{1}$$

where x is the vector of variables of the system. All variables, except for the interest rate and the interest spread, are measured in logarithms.[10] ε is a vector of serially uncorrelated reduced-form disturbances and $D(L)$ is a matrix polynomial in the lag operator L. The number of lags in the regression is determined by the Schwarz information criterion. According to this criterion, the optimal lag length for all systems is 2. The variance decomposition is computed by inverting the VAR to a vector moving average representation:

$$x_t = C(L)\varepsilon_t \tag{2}$$

and by orthogonalising the reduced-form residuals ε_t:

$$A_0 u_t = \varepsilon_t \tag{3}$$

where $E(u_t u_t') = I$. The orthogonalisation is done by a Cholesky decomposition, where the ordering corresponds to the ordering of the variables in the vector x. Thus, A_0 corresponds to the Cholesky decomposition of $\Omega = E(\varepsilon_t \varepsilon_t')$.[11]

In the following, I try to determine the information content of the financial variables for forecasting output and prices. To begin with, I consider the widely used 3-variable VAR consisting of logs of output y, prices p, and money m, so that the vector x corresponds to $x = [y\ p\ m]$.[12] In this study, money is represented by the aggregate M2. This aggregate is used by the Swiss National Bank as one indicator among others for predicting future price and output movements. The aggregate M2 is, however, not the intermediate target of the Swiss National Bank. Rather, the Bank sets a medium-term target for the monetary base. Since the demand for base money was hit by several structural shocks in the late 1980s, I prefer to use a broader aggregate in this study in order to examine the forecasting power of money on output and prices.[13] The $y\ p\ m$ VAR concentrates on the importance of money as a predictor of prices and output and does not consider any other financial variable. It therefore directly tests the monetarist hypothesis what movements in money are followed by subsequent movements in

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9 See, for example, Friedman and Kuttner (1996).

10 Small letters indicate variables in logs.

11 For the estimation of the VAR, a constant term is included.

12 For a discussion and survey of VAR studies, see e.g., Todd (1990).

13 See Rich (1997) for a complete analysis of the Swiss monetary policy in the post Bretton Woods period.

output and prices. However, it is important to note that the money component of the orthogonalised shocks (money innovations) of this VAR does not necessarily represent monetary policy shocks. Rather, the variance decomposition shows how important money innovations are for forecasting prices and output, independently of the true structural shocks, which cause unexpected changes in the variables of the system.

Table 1

Variance decomposition: *y p m* VAR

Variable	Horizon	Innovation		
		y	*p*	*M*
y	4	96 (1)	1 (0)	3 (1)
	8	93 (2)	1 (0)	6 (2)
	12	75 (5)	1 (1)	24 (5)
	16	62 (6)	1 (1)	37 (6)
	20	58 (7)	3 (1)	39 (7)
P	4	8 (4)	89 (4)	3 (1)
	8	23 (6)	70 (6)	7 (1)
	12	32 (7)	63 (7)	5 (1)
	16	34 (7)	60 (7)	6 (1)
	20	33 (7)	56 (7)	11 (1)

Note: In this and the following tables, standard errors (given in parentheses) were calculated using Runkle's (1987) bootstrapping method based on 200 replications.

Table 1 presents the variance decomposition for output and prices. All data used in this study consists of quarterly observations over the sample period from 1974:1 through 1996:4, so that the study covers the period of flexible exchange rates. The orthogonalisation of the system is made in the order of the vector and places output first, prices second, and money third. The standard errors are computed by using the bootstrap method by using the bootstrap method by Runkle (1987). Money innovations explain little of the forecast error variance up to a horizon of 8 quarters. However, for longer horizons, money becomes more important: At a 12-quarter horizon, money explains 25% of the output variance and at a 20-quarter horizon it explains almost 40%. This demonstrates clearly that money is an important predictor of real output, in spite of a substantial time lag between innovation and effect. The results are less favourable for the use of money as a useful information variable for prices. Money innovations explain very little of the forecast error variance of prices. Even at a 20-quarter horizon, only about 10% of the forecast error variance can be attributed to money innovations.

In the following, I expand the VAR analysis by including additional variables to find out whether other financial variables and asset prices contain information, which is not already included the monetary aggregate M2, to forecast prices and output. I am specifically interested in finding out whether the financial variable itself is important for the forecast error variance and whether the inclusion of the financial variable changes the relative forecasting power of M2. Therefore, I run a series of 4-variable VARs by including other financial variables, each of them in addition to M2. Thereby, I concentrate on 4 variables: the bond rate i (long-term interest rate), the trade-weighted nominal exchange rate index e, the SBC stock market index a, and the spread between the long and the short-run interest rate s.[14]

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[14] All data are from the Swiss National Bank data base. Output is measured by real GDP. Prices reflect the consumer price index. The monetary aggregate is M2. The bond rate corresponds to the long-term interest rate on government bonds. The

Table 2

Variance decomposition: *y p m i* VAR

Variable	Horizon	Innovation			
		y	*p*	*m*	*i*
Y	4	96 (1)	1 (0)	3 (1)	0 (0)
	8	92 (2)	1 (0)	5 (1)	2 (0)
	12	76 (5)	1 (1)	17 (4)	6 (1)
	16	66 (6)	1 (1)	25 (5)	8 (2)
	20	62 (6)	2 (1)	28 (6)	8 (2)
P	4	9 (5)	88 (5)	3 (1)	0 (0)
	8	24 (7)	70 (7)	6 (1)	0 (0)
	12	32 (8)	63 (8)	5 (1)	0 (0)
	16	35 (8)	59 (8)	5 (1)	1 (0)
	20	34 (8)	56 (8)	7 (2)	3 (1)

First, consider the *y p m i* VAR in Table 2. The inclusion of the bond rate *i* in the VAR slightly diminishes the forecasting importance of money for output. The bond rate itself is of little significance. When combined, *m* and *i* explain a smaller amount of the forecast error variance than *m* alone does in the *y p m* VAR. Of course, the ordering of the orthogonalisation favours *m* over *i* in the relative forecasting power. This is of some importance because of the relatively high correlation of the reduced-form residuals between *m* and *i*. However, *i* adds little new information to forecasting output. With respect to prices, the results are similar. The bond rate is unimportant for the forecast error variance of prices at all horizons considered.[15]

Table 3

Variance decomposition: *y p m e* VAR

Variable	Horizon	Innovation			
		y	*p*	*m*	*e*
Y	4	96 (1)	0 (0)	4 (1)	0 (0)
	8	92 (2)	1 (0)	4 (1)	3 (1)
	12	76 (4)	2 (1)	9 (3)	12 (3)
	16	65 (5)	2 (1)	15 (4)	18 (4)
	20	61 (6)	2 (1)	15 (4)	22 (5)
P	4	10 (4)	85 (4)	5 (1)	0 (0)
	8	30 (6)	48 (6)	19 (4)	3 (1)
	12	39 (6)	32 (6)	22 (4)	7 (2)
	16	40 (6)	26 (5)	19 (4)	15 (3)
	20	37 (6)	22 (5)	16 (3)	25 (5)

short-term rate is the 3-month interest rate. The exchange rate index is trade-weighted and nominal. The stock market index is computed by the Swiss Bank Corporation.

15 Similar results for both output and prices are obtained when the short-run interest rate is used instead of the long-run interest rate.

Second, I substitute the exchange rate index e for the interest rate i in the VAR. The results are reported in Table 3. The inclusion of e diminishes the forecasting importance of m for output. In addition, the exchange rate index e explains a substantial fraction of the output forecast error variance for horizons longer than 12 quarters. Furthermore, the inclusion of e drastically increases the significance of money for forecasting prices even at shorter horizons. The exchange rate index itself is also able to explain a large fraction of the forecast error variance of prices for horizons over 12 quarters. The financial variables m and e together are thus important information variables for forecasting prices.

Table 4

Variance decomposition: *y p m a* VAR

Variable	Horizon	Innovation			
		Y	*p*	*m*	*a*
y	4	97 (1)	0 (0)	2 (1)	1 (0)
	8	87 (3)	0 (0)	7 (2)	6 (2)
	12	67 (6)	1 (1)	23 (5)	9 (3)
	16	56 (7)	1 (1)	34 (6)	9 (3)
	20	53 (7)	2 (2)	37 (6)	8 (3)
P	4	11 (4)	85 (5)	1 (0)	3 (1)
	8	31 (7)	54 (8)	1 (1)	14 (4)
	12	38 (8)	38 (8)	1 (0)	23 (6)
	16	37 (8)	30 (7)	5 (2)	28 (6)
	20	33 (8)	26 (7)	11 (3)	30 (6)

Table 5

Variance decomposition: *y p m s* VAR

Variable	Horizon	Innovation			
		Y	*p*	*m*	*s*
y	4	97 (1)	0 (0)	3 (1)	0 (0)
	8	93 (2)	1 (0)	6 (2)	0 (0)
	12	77 (5)	1 (0)	22 (5)	0 (0)
	16	65 (7)	1 (1)	33 (7)	1 (0)
	20	61 (7)	2 (1)	36 (7)	1 (0)
P	4	7 (4)	90 (4)	2 (1)	1 (0)
	8	25 (6)	65 (6)	3 (1)	7 (2)
	12	32 (7)	54 (6)	2 (1)	12 (3)
	16	31 (7)	48 (6)	7 (1)	14 (3)
	20	28 (7)	44 (6)	15 (3)	13 (3)

Third, in place of the exchange rate index, the SBC stock market index a is included in the VAR. The results of the $y\ p\ m\ a$ VAR are shown in Table 4. The asset price index a itself explains only little of the output forecast error variance and leaves the fraction of the variance explained by money almost unchanged compared to the $y\ p\ m$ VAR. The inclusion of the stock market index does not improve the forecast power of money for prices. However, the stock market index alone seems to explain a large fraction of the forecast error variance of prices. Compared to the $y\ p\ m\ e$ VAR, m and a together explain less of the price forecast error variance than m and e.

Fourth, I consider the *y p m s* VAR, where the spread between the long and the short-run interest rate *s* is included (Table 5). With respect to output, the spread has almost no forecasting power at all. With respect to prices, the spread is more important than the bond rate, but less important than both the exchange rate index and the stock market index. Overall, the spread does not seem to be a very informative variable about future output and prices.[16]

Since the results indicate that *e* and *a* may be important, especially for forecasting prices, I run VARs which include either *e* or *a* but exclude *m*. The results of the *y p e* VAR are represented in Table 6. Compared to the *y p m* VAR, *e* explains approximately the same fraction of the output variance as does *m*. In addition, *e* has strong predictive content for prices for horizons beyond 10 quarters. On the contrary, *a* explains little of the forecast error variance for either output or prices (*y p a* VAR in Table 7). This indicates that the stock market index is of lesser importance for forecasting prices and output.

Table 6

Variance decomposition: *y p e* VAR

Variable	**Horizon**	**Innovation**		
		Y	***p***	***e***
Y	4	96 (1)	2 (1)	2 (1)
	8	80 (5)	5 (2)	15 (4)
	12	67 (7)	6 (3)	27 (6)
	16	59 (8)	6 (3)	35 (7)
	20	54 (9)	6 (3)	40 (8)
P	4	7 (5)	93 (5)	0 (0)
	8	30 (7)	63 (7)	7 (2)
	12	42 (6)	37 (6)	21 (4)
	16	42 (6)	24 (4)	34 (6)
	20	39 (7)	18 (3)	44 (7)

Table 7

Variance decomposition: *y p a* VAR

Variable	**Horizon**	**Innovation**		
		Y	***p***	***e***
y	4	100 (0)	0 (0)	0 (0)
	8	99 (0)	1 (0)	0 (0)
	12	98 (0)	2 (0)	0 (0)
	16	97 (1)	2 (1)	1 (0)
	20	95 (1)	4 (1)	1 (0)
P	4	20 (6)	77 (7)	3 (1)
	8	49 (8)	44 (8)	7 (2)
	12	63 (7)	29 (7)	8 (2)
	16	70 (7)	22 (6)	8 (2)
	20	74 (7)	19 (5)	7 (2)

[16] Note that the spread is I(0). Thus, by estimating the VAR in levels, the information content of the spread may be underestimated.

The results from these variance decompositions lead to the conclusion that both money and the exchange rate index are important information variables for forecasting output and prices. They dominate the other financial variables, i.e., the bond rate, the spread and the stock market index as predictors of output and prices at all horizons. Money and the exchange rate index are especially helpful for forecasting at horizons over 8 to 10 quarters. For forecasts up to 8 quarters, the time series of output and prices seem to contain the most predictive information themselves.

2. Out-of-sample forecasts

Variance decompositions reflect the in-sample fit of vectorautoregressions and thus concern the in-sample forecasting ability. However, superior in-sample forecasting ability does not automatically mean superior out-of-sample forecasting ability. As put forward by Bernanke (1990) and by Thoma and Gray (1994), the ultimate decision about the usefulness of a variable as an information variable must come from its ability to forecast out of sample. Consequently, I test the out-of-sample forecasting ability with respect to output and prices of different VAR systems by applying a variation of the method used in Thoma and Gray (1994). The out-of-sample forecasting ability is measured by the root mean square error of forecasts at different horizons. The statistic for the 4-quarter horizon is computed as follows: The VAR is estimated over the period 1974:3–1984:2 (40 observations). Using the estimated coefficients and dynamic forecasting techniques, forecasts of output and prices in 1985:2 are generated (4-quarter-ahead forecasts). Then, the sample is shifted one quarter ahead to cover the period 1974:4–1984:3. The VAR is now re-estimated, so that forecasts for 1985:3 can be generated. This procedure is continued until the forecasts reach the end of the sample (1996:4). The generated series of 4-quarter-ahead forecasts and the actual data can be used in order to compute the root mean squared forecast error. The same method is applied to compute 8 and 12-quarter horizon forecasts and the corresponding root mean square errors (RMSE). The RMSE of the different VAR systems can then be used to evaluate forecasting ability. The smaller the RMSE, the more information is contained in the variables of the VAR considered.

Table 8

RMSE of 4-quarter-ahead forecasts

VAR system	RMSE for y	RMSE for p
$y\ p$	0.0265	0.0206
$Y\ p\ m$	0.0180	0.0197
$y\ p\ m\ I$	0.0205	0.0213
$y\ p\ m\ e$	0.0175	0.0165
$y\ p\ m\ a$	0.0199	0.0240
$y\ p\ m\ s$	0.0181	0.0197
$y\ p\ e$	0.0276	0.0194
$y\ p\ a$	0.0350	0.0184

Table 8 reports the results for the 4-quarter forecasting horizon. I consider all the VAR systems which were analysed for the in-sample forecasting ability. Furthermore, I include the two variable $y\ p$ VAR in the analysis. This VAR can be used as a direct benchmark in order to find out whether the inclusion of a specific information variable helps to reduce the RMSE for output and prices. Such a comparison was not possible in the analysis of the variance decompositions of Section 1.

The inclusion of M2 in the VAR system ($y\ p\ m$ VAR) helps to reduce the RMSE of output by almost a third, but the RMSE of prices is only changed to a small extent. Thus, money contains useful information about output over the next 4 quarters. With respect to the forecasts of prices, the information content of M2 is small. Next, I increase the VAR system to contain 4 variables by adding each time another financial variable in addition to M2. The *y p m i* VAR brings no improvement over the *y p m* VAR. Thus, the bond rate does not dominate M2 as an information variable over the 4-quarter horizon. The results are similar for the *y p m a* VAR and the *y p m s* VAR. Whereas the *y p m a* VAR worsens the results, the *y p m s* VAR achieves almost the same RMSEs as the *y p m* VAR. On the contrary, the *y p m e* VAR improves the forecasts for both output and inflation. The decline in the RMSE for output is only small, whereas for prices the reduction is quite large (approximately 15%). Thus, the exchange rate index again seems to be an important information variable, especially for prices. In order to find out whether the exchange rate index contains forecast information independently of M2, I compute the RMSEs from the *y p e* VAR, where M2 is dropped from the system. The RMSE for both output and prices becomes larger, indicating that the exchange rate index delivers the best forecasting information (for the 4-quarter horizon) if combined with a monetary aggregate.[17]

Table 9

RMSE of 8-quarter-ahead forecasts

VAR system	**RMSE for *y***	**RMSE for *p***
y p	0.0564	0.0402
Y p m	0.0275	0.0429
y p m I	0.0330	0.0499
y p m e	0.0302	0.0384
y p m a	0.0320	0.0515
y p m s	0.0304	0.0425
y p e	0.0563	0.0336
y p a	0.0696	0.0484

Table 10

RMSE of 12-quarter-ahead forecasts

VAR system	**RMSE for *y***	**RMSE for *p***
y p	0.0863	0.0604
y p m	0.0439	0.0656
y p m I	0.0455	0.0753
y p m e	0.0510	0.0593
y p m a	0.0507	0.0783
y p m s	0.0520	0.0633
y p e	0.0867	0.0461
y p a	0.0980	0.0854

Table 9 shows the findings for the 8-quarter horizon from the same VAR systems. The *y p m* VAR performs best with respect to output. The *y p m* VAR cuts the RMSE for output in half compared to the *y p* VAR. Furthermore, all 4-variable systems have bigger RMSEs for *y* than the *y p m* VAR. As for the 4-quarter horizon, the inclusion the exchange rate index (*y p m e* VAR) improves the forecast for prices. However, the best result is achieved if M2 is dropped from the VAR, which confirms the importance of exchange rates for forecasting prices.

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[17] As a reference, the results from the *y p a* VAR are also included in Tables 8 to 10.

In Table 10, the results for the 12-quarter horizon are presented. They are similar and consistent with the findings for the other horizons: The *y p m* VAR performs best with respect to output and the *y p e* VAR achieves the best out-of-sample forecasts for prices. The out-of-sample forecast analysis shows that money (M2) and the exchange rate index *e* are the two most important information variables of all the financial variables considered. Interest rates, interest rate spreads, and stock market prices do not seem to incorporate superior information, which is not already reflected by either money or exchange rates. The results of this section are consistent with those obtained from the variance decomposition. In both experiments, money and the exchange rate turned out to be the most important information variables. However, the variance decompositions indicated that these variables are only of interest for medium and long-term forecasts, whereas the out-of-sample forecasting exercise shows that money and exchange rates are also important for short horizon forecasts (e.g., over 4 quarters). The exercise carried out in this section also shows that expanding the VAR to include more variables does not generally improve the forecasting ability and may actually cause a decline in the out-of-sample forecasting power. This confirms the findings by Thoma and Gray (1994).

3. The change of the predictive content

The information content of the variables may change over time because of structural shocks to the economy. So far, the analysis did not take up this problem. In this section, I check whether the forecasting power of the variables changes over time. Thereby, I concentrate on the 3 systems, which proved to be most valuable for forecasting prices and output, namely on the 3-variable VARs *y p m* and *y p e* and the 4-variable VAR *y p m e*. I use a rolling regression methodology explained in detail below. This kind of technique was used, for example, in Thoma and Gray (1994), Friedman (1996), and Friedman and Kuttner (1996).

I start by considering the in-sample forecasting ability and compute variance decompositions for the 3 systems. I chose a series of consecutive sample periods each consisting of 40 observations (10 years). The first sample starts in 1974:3 and ends in 1984:2. For each sample, the fraction of the forecast error variance is computed for both output and prices over the 8 and the 12-quarter horizons. The 4-quarter horizon is not reported, because the fraction of the forecast error variance explained by financial variables is generally very small. The percentages of the forecast error variance due to financial variables are shown in Figures 1 to 3. The horizontal line indicates the last observation of the estimation sample.

The results from the *y p m* VAR are plotted in Figure 1. Generally, the importance of money for the forecast error variance is very sensitive to the sample period. Shifting the end of the sample towards 1989 increases the importance of M2 strongly. Up to 60% of the forecast error variance of output is explained by money even at the 8-quarter horizon. However, if the end of the sample is expanded beyond 1994, money explains only a small fraction of the forecast error variance, indicating that the predictive content of money has become smaller. The results are similar for prices. Shifting the sample forward to 1992 sharply increases the fraction of the forecast error variance explained by M2. Shifting the sample beyond 1993 causes a deterioration of the forecasting ability.

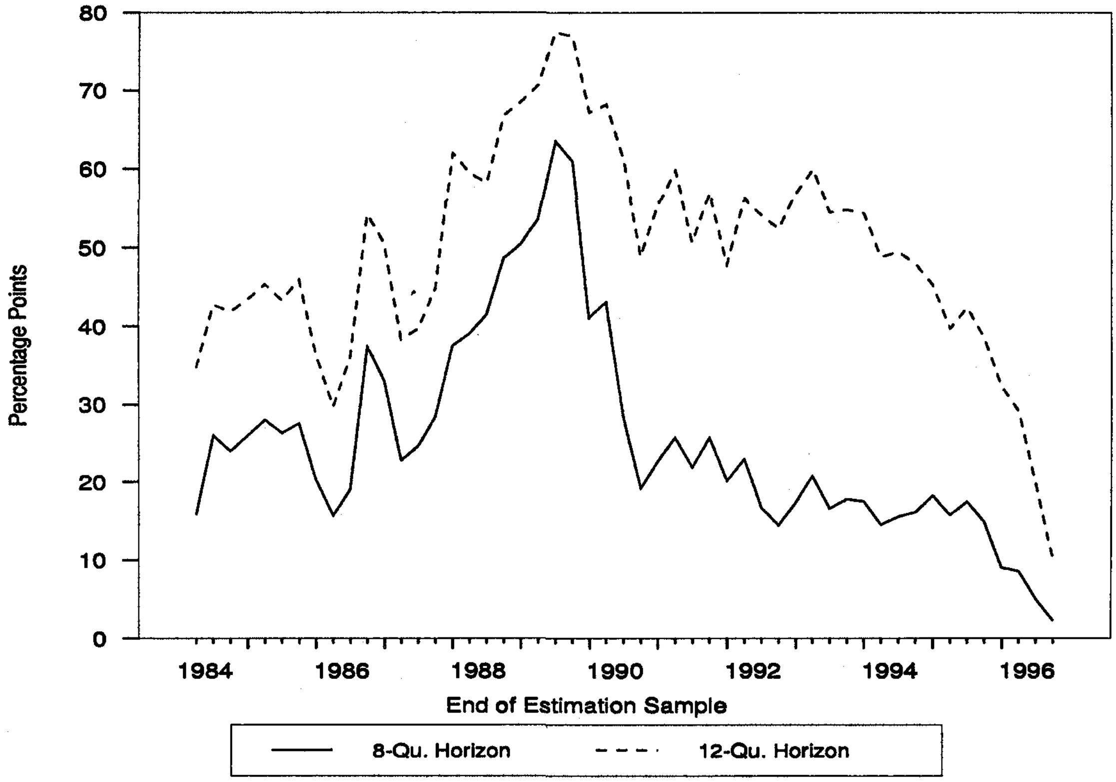
In Figure 2, the results from the *y p e* VAR are shown. It can be observed that the exchange rate index contains important information for samples ending before 1986. In these samples, large fractions of the forecast error variance of both output and prices are due to innovations in the exchange rate index. However, the predictive content of the exchange rate index sharply deteriorates if the sample ends after 1986.

Figure 3 presents the findings for the *y p m e* VAR. The most striking result is that the information content of money is deteriorating to a lesser extent for samples ending after 1989 if the exchange rate is included in the system. The importance of the exchange rate itself generally declines if the sample is shifted forward. However, the forecasting ability of money is much more robust if the

Figure 1

***y p m* VAR**

A: Forecast error variance of output explained by money



B: Forecast error variance of prices explained by money

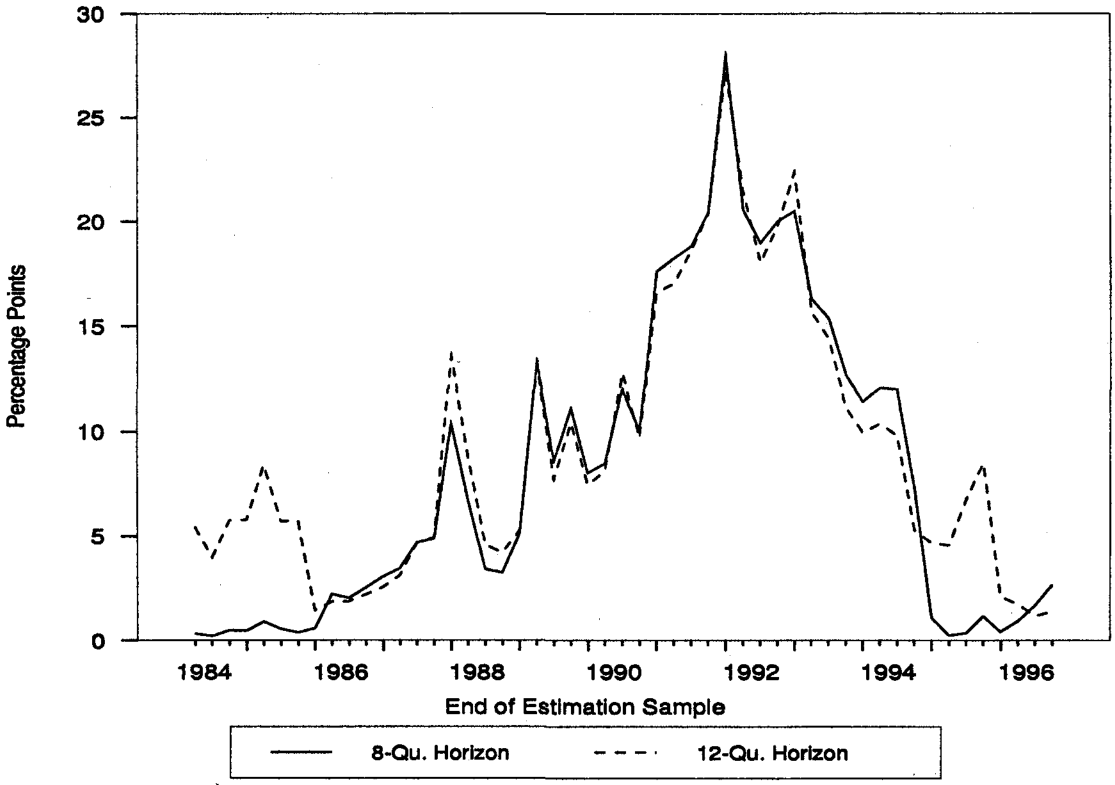
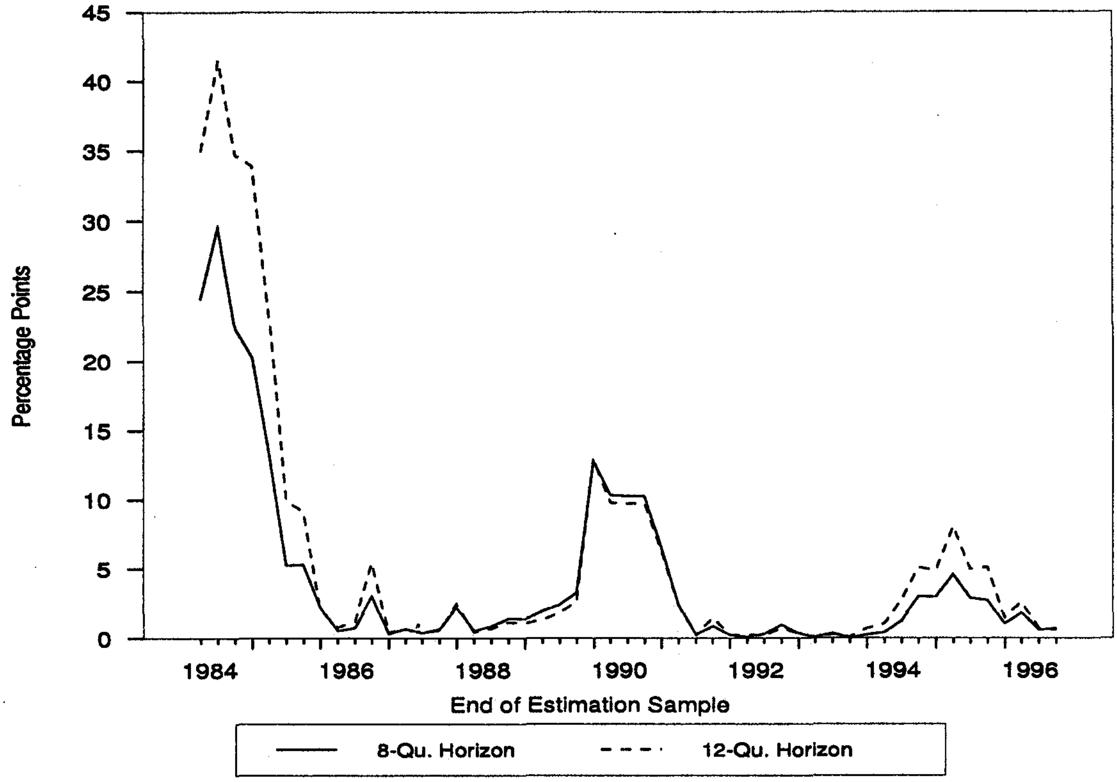


Figure 2

y p e VAR

A: Forecast error variance of output explained by exchange rates



B: Forecast error variance of prices explained by exchange rates

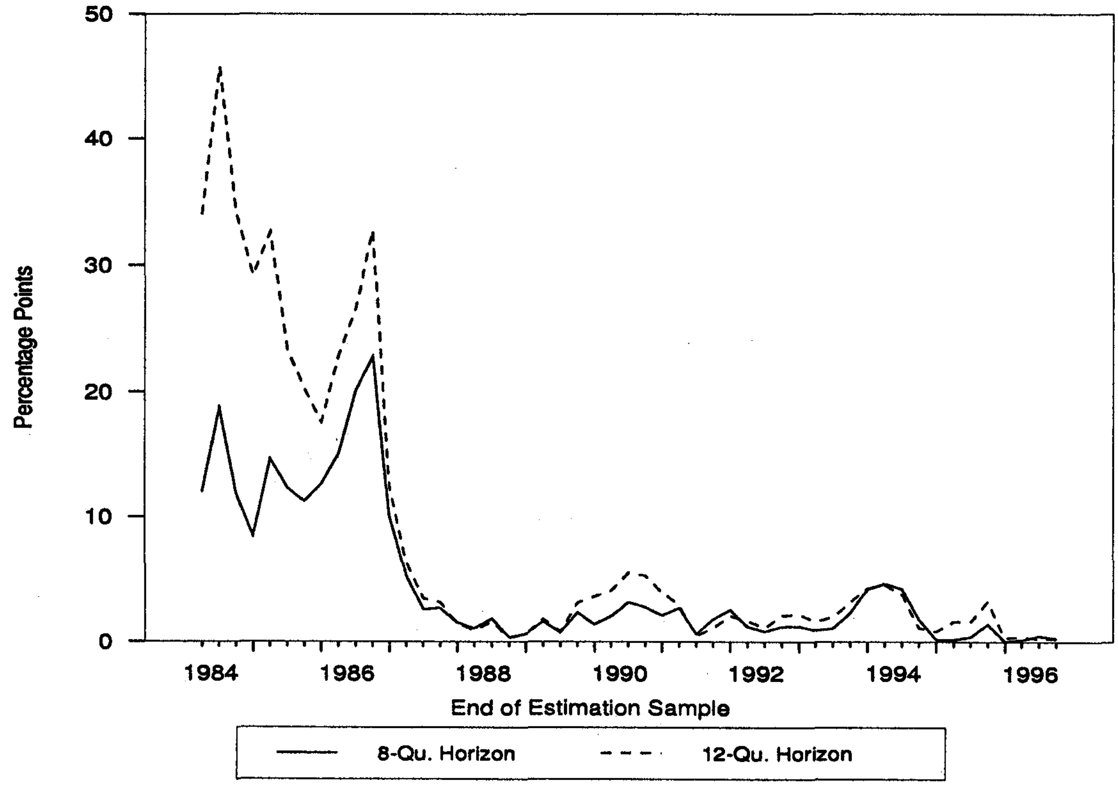


Figure 3

y p m e **VAR**

A: Forecast error variance of output explained by money

B: Forecast error variance of output explained by exchange rates

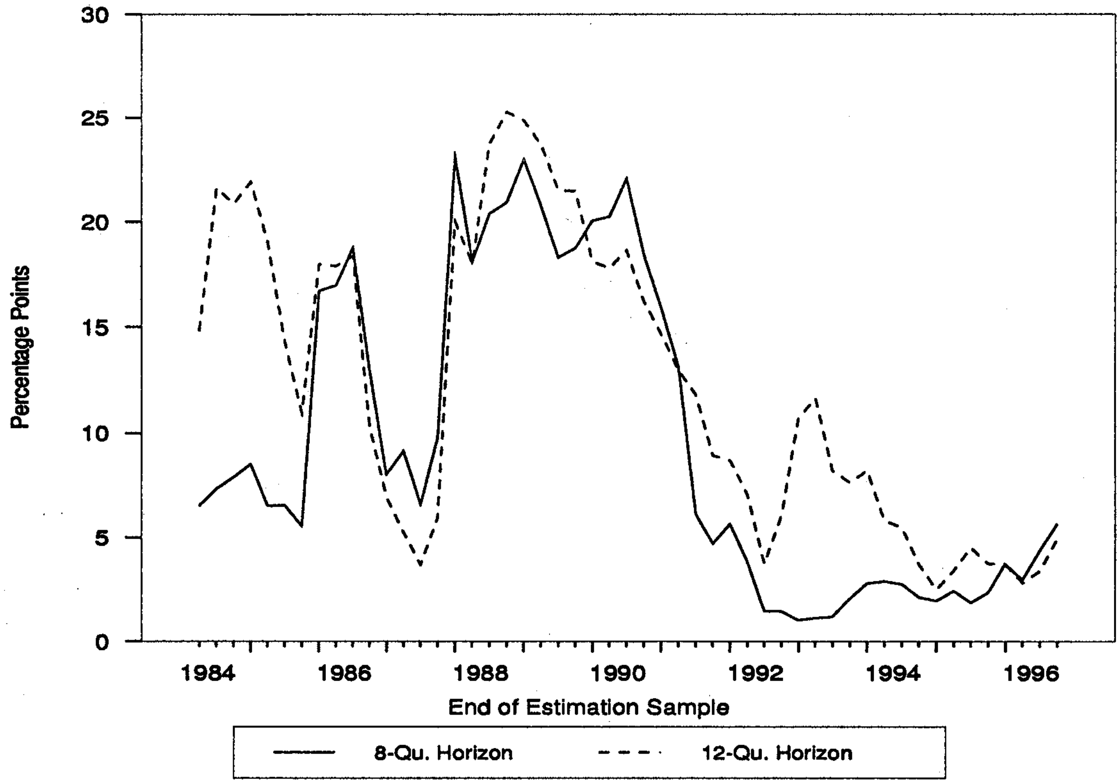
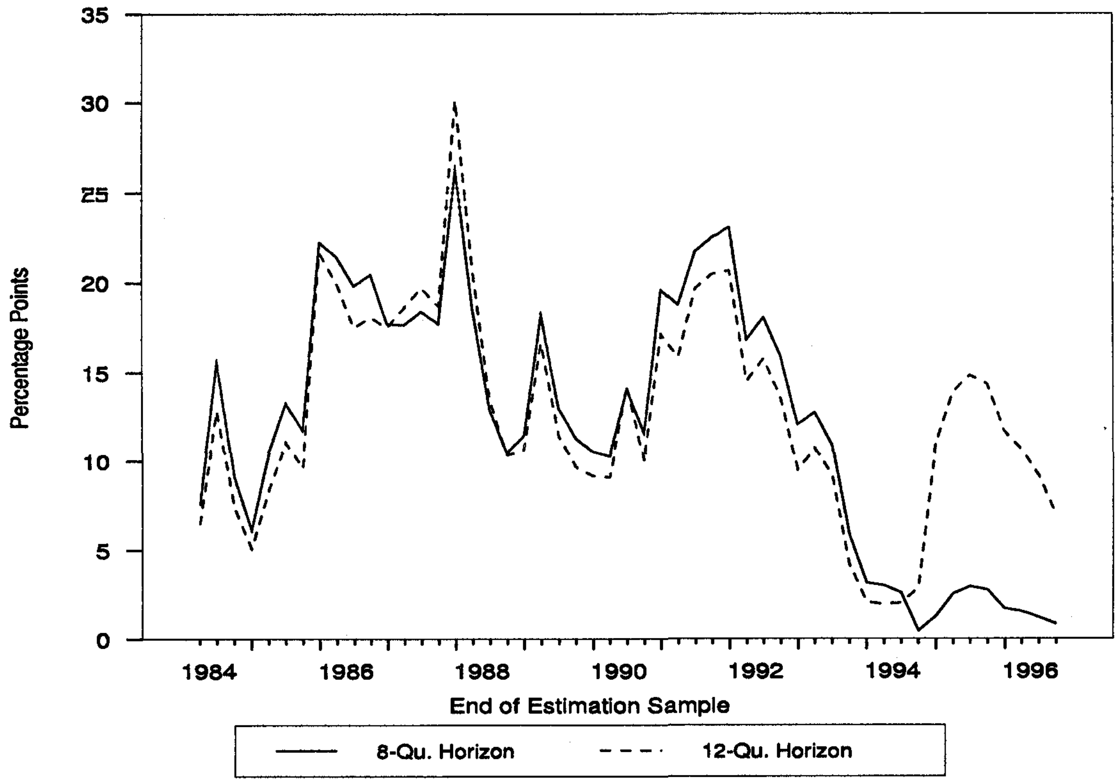


Figure 3 (cont.)

y p e **VAR**

C: Forecast error variance of prices explained by money



D: Forecast error variance of prices explained by exchange rates

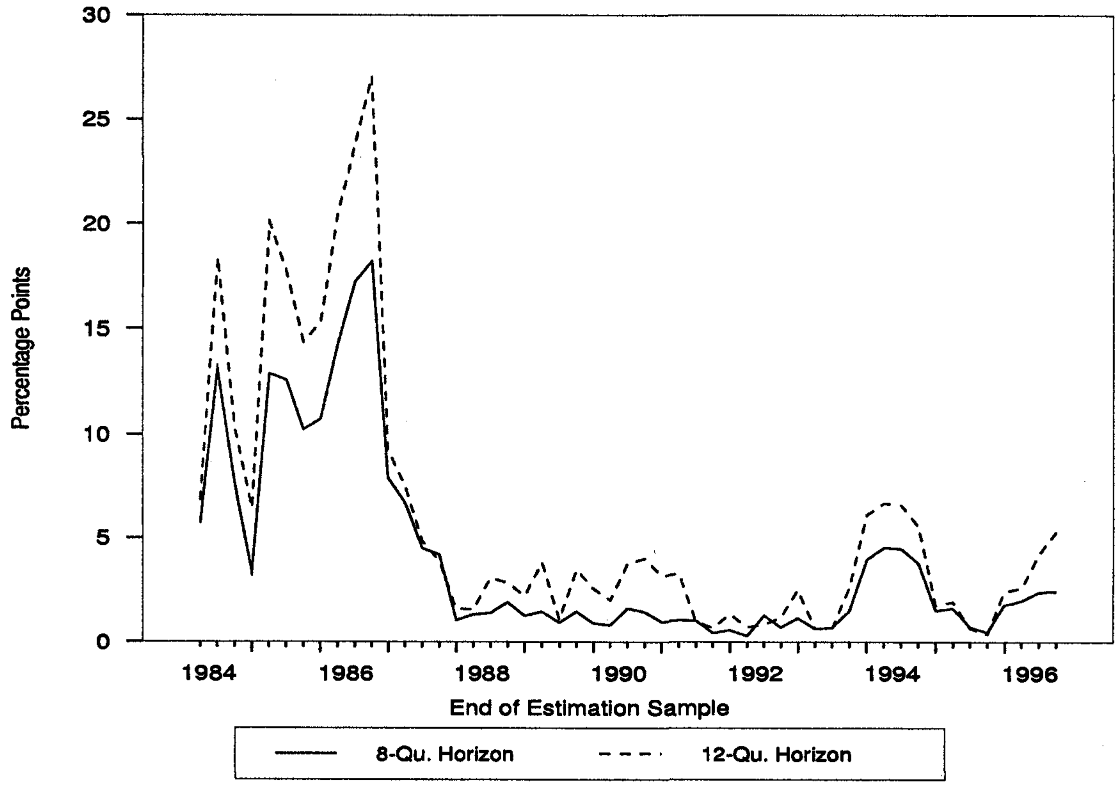
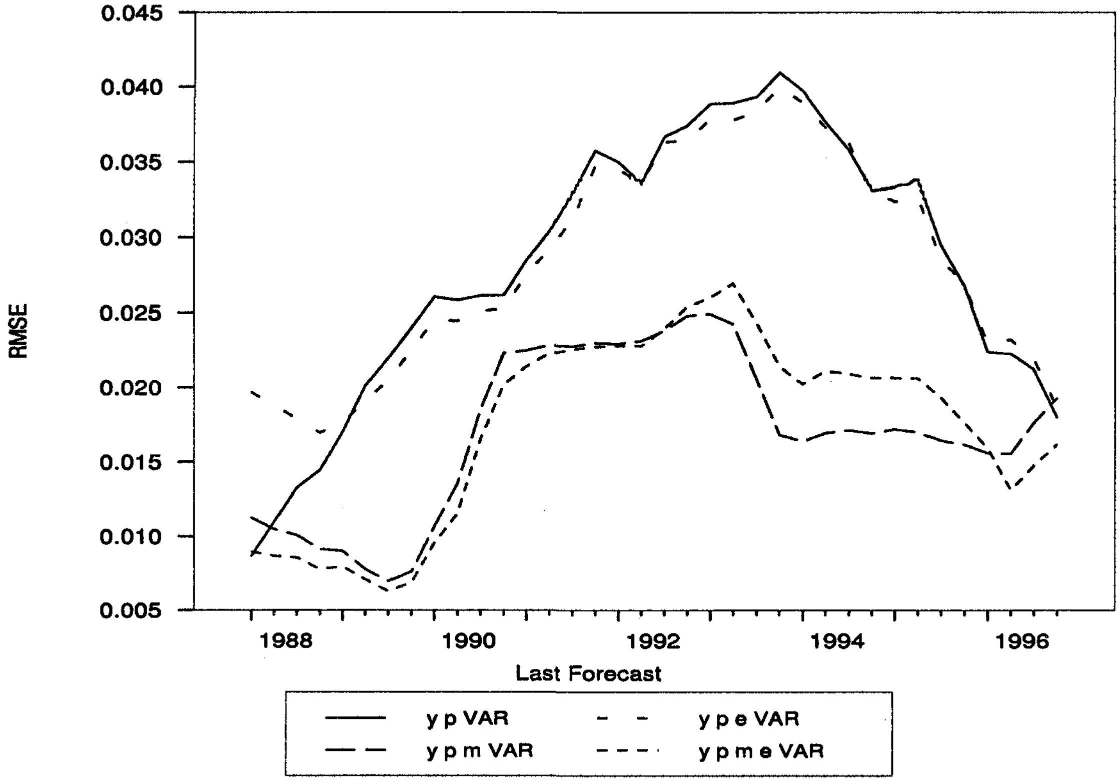


Figure 4

Root mean square error of 4-quarter-ahead forecasts

A: Output



B: Prices

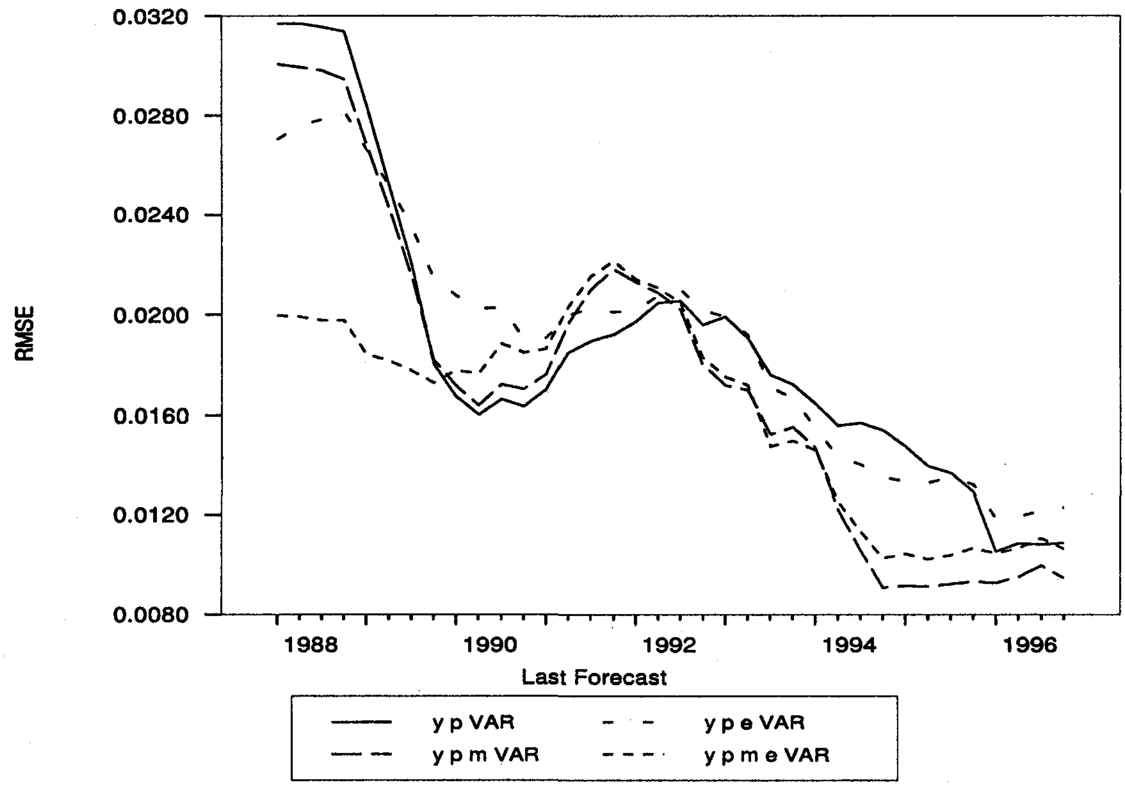
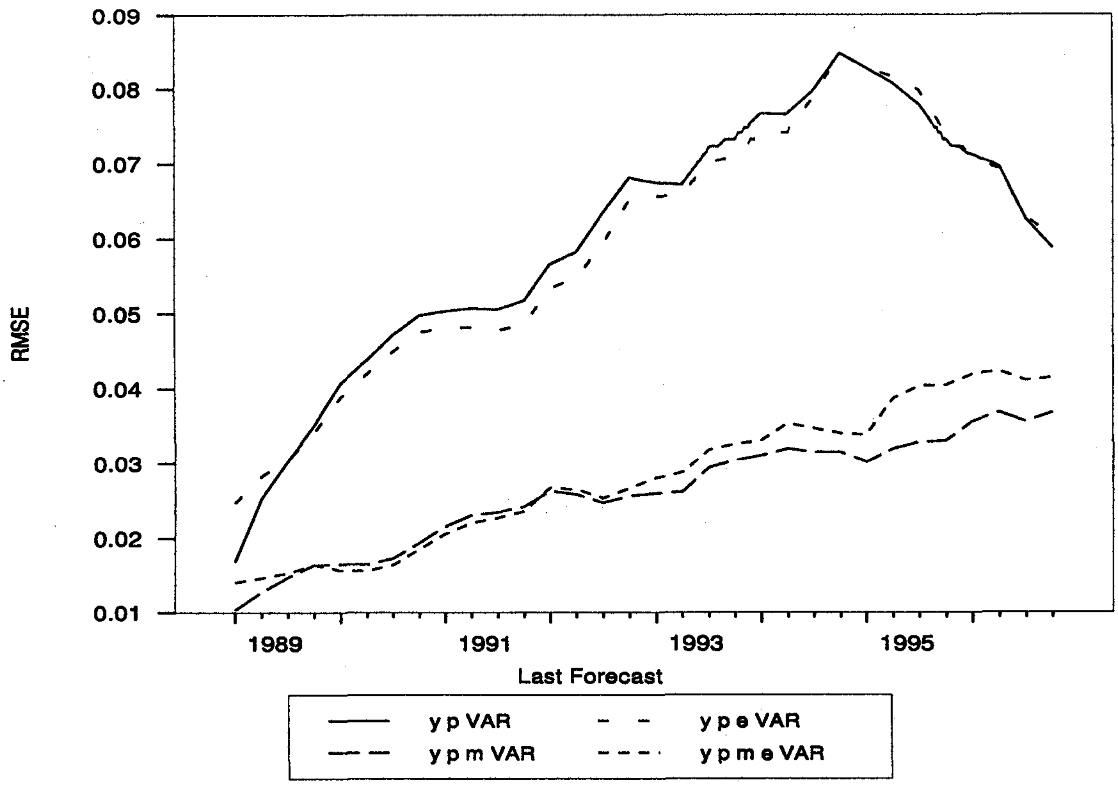


Figure 5

Root mean square error of 8-quarter-ahead forecasts

A: Output



B: Prices

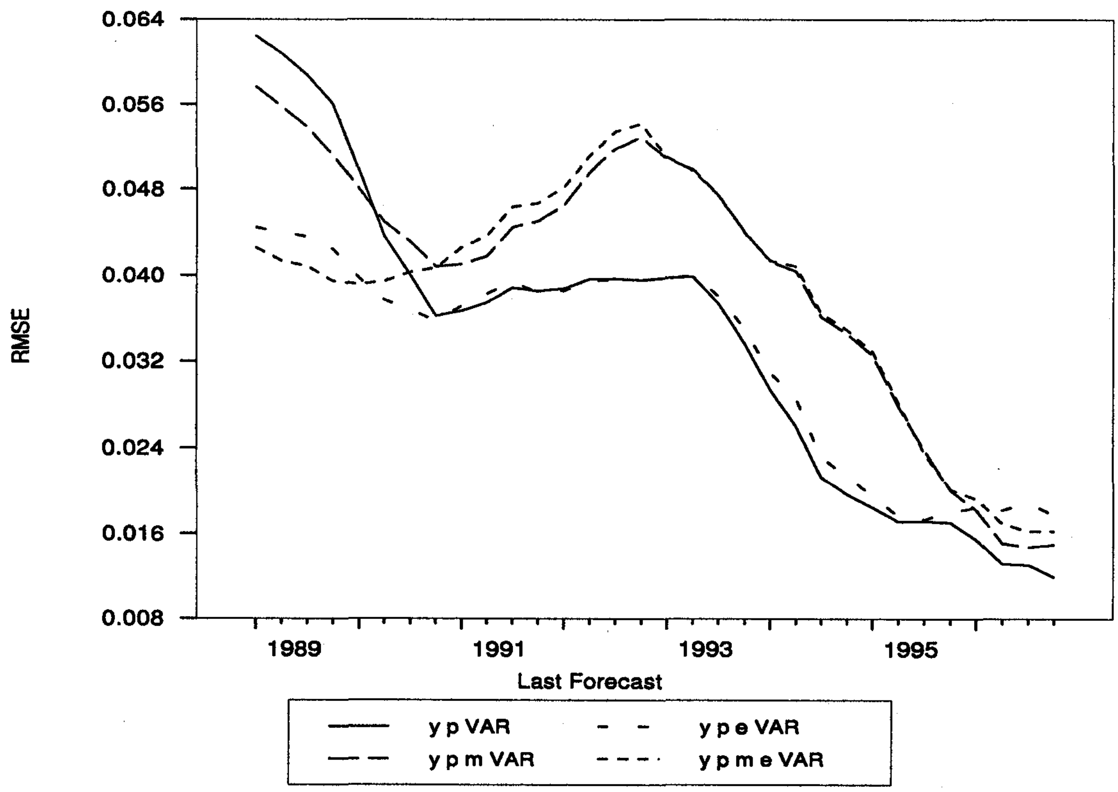
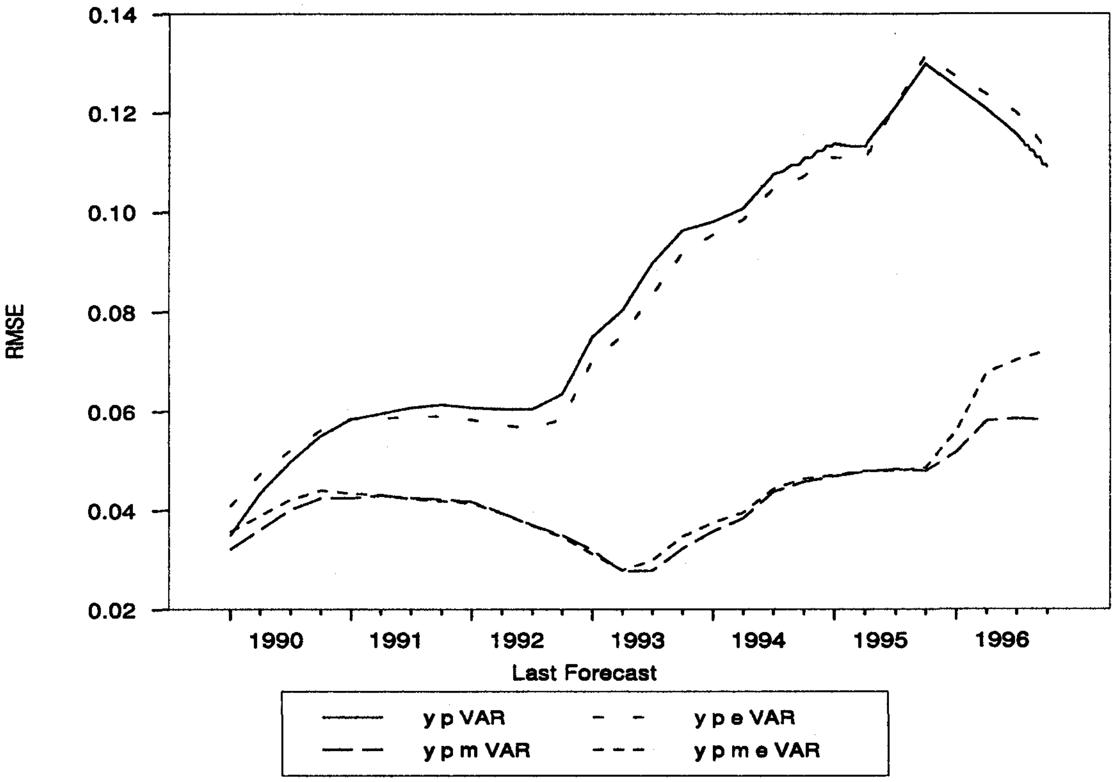


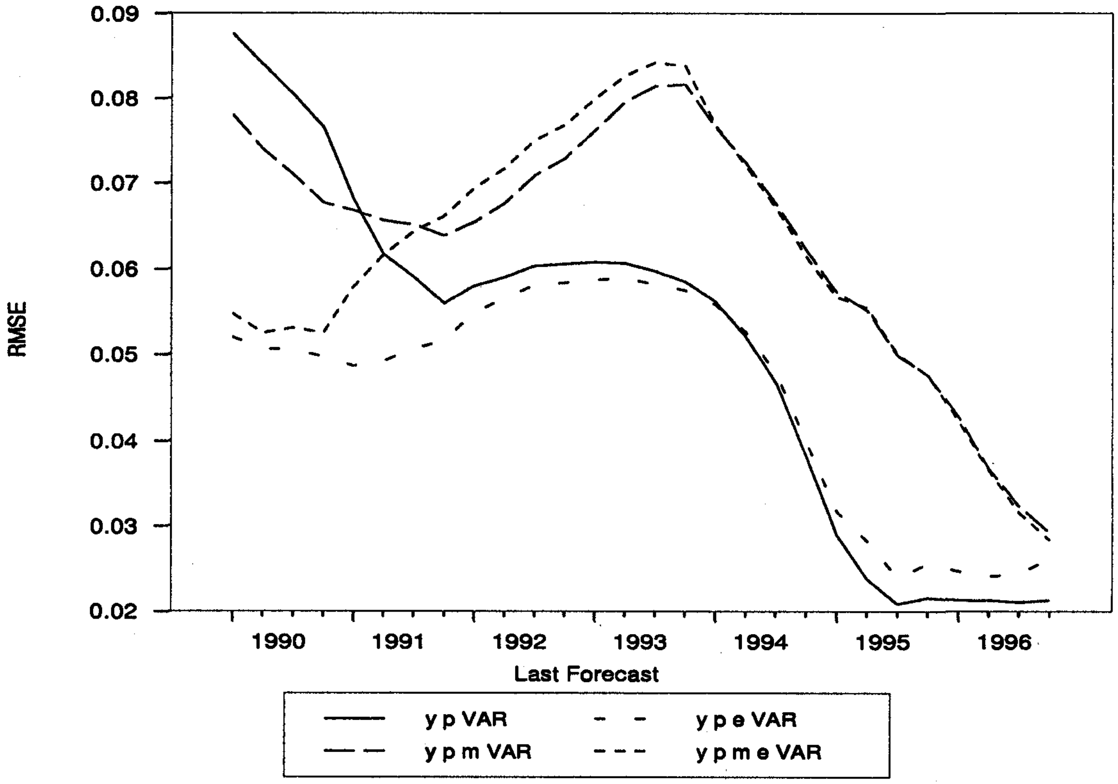
Figure 6

Root mean square error of 12-quarter-ahead forecasts

A: Output



B: Prices



exchange rate index is included in the VAR system. This underlines that money and exchange rates together contain more stable information about prices and output than each of these variables considered alone.

Next, I consider the evolution of the out-of-sample forecasting ability. I compute forecasts for the 4, 8 and 12-quarter horizon as explained in Section 2 based on samples of 40 observations. Then, the root mean square error is computed for 12 consecutive forecast errors. For example, the RMSE reported in 1988:1 refers to the 12 forecast errors from 1985:2 through 1988:1 and the RMSE reported in 1988:2 refers to the 12 forecast errors from 1985:3 through 1988:2 and so on. This is done for all 3 forecast horizons. In order to find out whether the financial variables keep their predictive information content for output and prices, I also include the $y\ p$ VAR in the analysis.

The results are reported in Figures 4 to 6. In general, the RMSE of output becomes bigger, the more recent the data included in the sample, but the RMSE for output of both the $y\ p$ VAR and the $y\ p\ e$ VAR improves again if data after 1994 are included. Conversely, the RMSE for prices generally declines if the sample is shifted forward. This is true for all 4 different VARs considered. There are two important findings: First, the information content of the financial variables, especially the exchange rate index, for forecasting prices deteriorates over time. The simple $y\ p$ VAR achieves the lowest RMSE at the end of the observed data, compared with the larger systems with the exception of the 4-quarter horizon, where all systems achieve similar RMSEs. Second, for forecasting output, the inclusion of money still improves the forecasts at longer horizons, i.e., the systems including m achieve lower RMSEs than those without money. This is less clear for the 4-quarter horizon. These findings confirm that forecasting output has recently become more difficult relative to forecasting prices. Whereas money and exchange rates may still improve output forecasts, they now contain little information about prices at the horizons considered. The out-of-sample results are basically consistent with the in-sample findings. The only discrepancy lies in the different judgements about the information content of the exchange rate index for predicting output.

Conclusions

This study reconsidered the information content of a set of financial variables for forecasting output and prices in Switzerland during the post Bretton Woods era. The analysis covers three parts. First, the in-sample predictive content of the variables is analysed. Second, the out-of-sample forecasting ability is considered. Third, the study asks whether the information content changes over time. In all parts of the paper vectorautoregression methodology is applied, so that the information content of the financial variables is measured as the additional predictive information which is not already extractable from observing the time series of output and prices themselves.

The results show that money contains important information for forecasting output. The information content has recently declined, but forecasts based on systems including money still outperform systems without money. Including the exchange rate index in the forecast system may render the forecast ability more stable over time, but the evidence is not clear-cut. Exchange rates turned out to be helpful for predicting prices. However, the information content of the exchange rate index has deteriorated strongly in recent years, so that the best forecasts for prices are based on a forecasting system including only output and prices.

The results from this exercise lead to three conclusions. 1. Financial variables and asset prices lose information content for predicting output and prices during the 1990s. 2. Forecasting prices gradually becomes easier during the 1990s because more information is contained in the time series of prices itself. This indicates that the inflation process may have changed during the 1990s. In contrast, forecasting output becomes gradually more difficult especially at longer horizons because of the loss of information of money. 3. The information contents of the VAR systems are not robust over time. Rolling regression techniques may help to find out whether one VAR system gradually becomes less attractive than another for forecasting output and prices.

It is important to remember that the information content of the financial variables is defined in a specific statistical manner. The information content of a specific variable is measured as the incremental predictive power over the part of movements of output and prices that is not already forecastable from past values of output, prices and, in general, from the variables placed ahead in the order of orthogonalisation. If monetary policy consists of systematic responses to past fluctuations in output and prices, the information content of money may be small although its impact on output and prices may be large. Furthermore, if money growth is relatively stable, the correlation between money and prices and between money and output can be small even if money has powerful effects. Consequently, knowing that the information content of money is small for prices and declining for output does not mean that the central bank should abandon monetary aggregates as information variables or as intermediate targets. In addition, none of the other financial variables considered delivered better information for forecasting prices and output in a consistent manner. Thus, the study does not favour the use of such variables as indicators for monetary policy in Switzerland and underlines the difficulties the Swiss National Bank would face if it pursued a policy of inflation targeting instead of a policy of monetary targeting.

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How informative are financial asset prices in Spain?

Francisco Alonso, Juan Ayuso and Jorge Martínez-Pagés

Introduction

Agents participating in financial markets are often characterised as being forward-looking. Accordingly, financial prices can be considered forward-looking regarding those macroeconomic variables that affect them and, therefore, should contain valuable information on their future or expected behaviour. Moreover, in comparison with other potential sources of information, financial prices are easier and cheaper to obtain and can be recorded for higher frequencies.

Unsurprisingly, then, there is a relatively extensive literature focused on extracting the informational content of financial prices on future macroeconomic fundamentals. In the early 1990s a number of papers analysed the US case and found that several financial indicators, mainly those related to the term structure, provided reliable information on future interest rates (Campbell and Shiller (1991)), inflation (Mishkin (1990)) or real activity (Estrella and Hardouvelis (1991)). Similar results were later found for other economies (Estrella and Miskhin (1996), Davis and Fagan (1996), Bernard and Gerlach (1996)).

This paper builds on this literature and attempts to analyse the informational content of financial prices in Spain, mainly from the viewpoint of a central bank. There are two main reasons why this analysis of the Spanish case may be relevant. First, the process of liberalisation and modernisation of the Spanish financial system, though extraordinarily fast, was initiated only very recently compared to other Western countries. Indeed, until very recently, there were no data covering a period long enough to allow a systematic analysis of the informational content of financial indicators. Even now, data are still insufficient or of poor quality in some cases. This explains why the issue has not been studied much in Spain.[1]

Second, until 1994 Spanish monetary policy followed a classical two-level strategy, with a monetary aggregate playing the role of an intermediate target. In this framework, monetary indicators pushed other indicators to a secondary level of importance. Since 1995, a new monetary strategy has been implemented in which inflation is directly targeted. This new framework has provided scope for other non-monetary indicators, among which financial indicators are potentially useful. In particular, there is a new demand for indicators in order to make projections regarding relevant macroeconomic variables. Those variables are typically inflation, short-term interest rates and also output. As recently stressed in Svensson (1997), direct inflation targeting does not necessarily imply that a central bank should not worry about output deviations from a reference or targeted level.

This paper examines, from an empirical standpoint, the informational content of the financial indicators most commonly considered in the literature: domestic yields and yield spreads, foreign-domestic spreads, credit quality spreads, stock prices and exchange rates. We focus on their informational content with respect to the inflation rate, the 3-month interest rate and output.

As to the methodology, since we aim to provide an overall view of the usefulness of these indicators, we consider three alternative approaches. First, we analyse the predictive power of financial prices by comparing the out-of-sample performance of equations containing each financial indicator with a simple univariate equation containing only lagged values of the dependent variable. Next, following a recent work by Estrella and Miskhin (1996), we also address the possibility of using

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1 Some exceptions are Martínez-Resano (1993), Davis and Fagan (1996) or Alonso et al. (1997).

financial prices as "qualitative" indicators and estimate Probit models to forecast inflationary upturns, output slowdowns and monetary policy tightenings as reflected by interest rate upturns. Finally, we analyse the possibility of using financial prices as expectation indicators, independently of their ability as predictors.

The structure of the paper is as follows. Section 1 presents our methodological approach to assessing the quantitative informational content of the different indicators considered. The main results of this approach are presented in Section 2, showing that, in general, financial prices do not outperform simple univariate models. Given this result, two alternative routes are further explored. In Section 3, the results of a rough approach to analysing the usefulness of financial prices as "qualitative" indicators to predict specific episodes are presented. In Section 4 we comment on the relationship between predictors and expectation indicators and consider the potential usefulness of financial prices as indicators of expectations of future inflation and interest rates. The final section summarises the main conclusions of the analysis and extracts some policy implications.

1. An approach to assessing the quantitative informational content of Spanish financial prices

1.1 Empirical strategy

It is not an easy task to come to any conclusion on the informational content of a variable regarding the future behaviour of another. Such an assessment will always be conditional upon, at least, three assumptions: first, the information set included (the indicator, the indicator plus lagged values of the variable to be forecast, third variables, etc.); second, the predictive horizon we are interested in; and third, the criterion for assessing performance. Before presenting our approach, it is worth reviewing the competing alternatives to specifying the relevant assumptions.

Most papers in the existing literature follow what we could call a "basic approach": one or several regressions are run in which the macrofundamental to be predicted is on the left-hand side and (some transformation of) the indicator is included on the right-hand side. Apart from this common root, differences are considerable. Regarding the specification of the information set, some authors take a static bivariate approach in which the indicator, usually lagged, is the only regressor (Mishkin (1990)). Others also use a bivariate model but follow a "Granger causality" approach, thus introducing some dynamics in the analysis and considering lagged values of both the dependent variable and the indicator on the right-hand side (Davis and Fagan (1996)). A third approach consists of including on the right-hand side of the equations several indicators to allow for some competition among them (Bernanke (1990)). Finally, there are also examples of VAR analysis in which more than one fundamental is predicted simultaneously (Davis and Fagan (1996)).

Regarding horizons, most papers consider several horizons simultaneously, with special attention paid to the distinction between the short and the long term. As to the performance criterion, two main approaches can be mentioned. In some papers, usual goodness-of-fit in-sample statistics are used to test the significance of the indicators in the regressions and their contribution to reducing the residual standard error. Other papers, however, focus on the out-of-sample forecasts.

Our aim in this paper is to analyse to what extent financial prices contain useful information for the Spanish monetary authorities on the future or expected behaviour of inflation, output and short-term interest rates, other than the information that the past pattern of each macroeconomic variable can provide. Thus, we will consider equations including lagged values of the dependent variable and lagged values of the financial indicators. In particular, we consider up to 12 quarterly lags which provide a maximum delay of 3 years between the indicator and the fundamental.

Nevertheless, we do not combine either macrofundamentals or indicators. Our data base does not cover a period long enough to allow a more complex analysis in which we could look at more than one indicator – or more than one fundamental – at the same time.

Regarding the performance criteria, although we test in-sample joint significance we focus on out-of-sample properties to assess the usefulness of the different indicators. In particular, we compare the mean squared errors of forecasts 1, 4, 8 and 12 quarters ahead of both the univariate equation and the equation including the indicator. Therefore, the prediction horizons span 1 quarter to 3 years.

Our approach can be summarised in the following steps:

1. A univariate autoregressive model is estimated for quarterly data on the (stationary transformation of the) macrofundamental y:

$$y_t = a_0 + \sum_{i=1}^{p} a_i y_{t-i} + \varepsilon_t \tag{1}$$

The maximum lag p has been chosen testing the estimated residual autocorrelations, the joint significance of the included lags and the joint (non-)significance of the excluded lags between 1 and 12.

2. We check the order of integration of the indicator. If the macrofundamental and the indicator are of the same order, we check whether they are cointegrated.[2] If this is the case, a lagged standard error correction term, *ecm*, and 12 lagged values of the (stationary transformation of the) indicator x are added. If there is no cointegration, only the 12 lags are included. In both cases, the joint significance of the new regressors is tested. If they are not significant, we stop the analysis and conclude that this is not a useful indicator. If they are significant, the following exercise is undertaken to determine the length of the lag polynomial: the first and/or last lags are subsequently excluded and, after each exclusion, the joint significance of the included lags and the joint (non-)significance of the excluded ones is tested. This yields the following equation:

$$y_t = a_0 + \sum_{i=1}^{p} a_i y_{t-i} + \sum_{j=q1}^{q2} b_i x_{t-j} + \delta_x c_1 ecm_{t-1} + v_t \tag{2}$$

where $q1 \geq 1$, $q2 \leq 12$, and δ_x is equal to 1 if there is cointegration between the fundamental and the indicator, and 0 otherwise. Notice that the same number of lags (p) for the dependent variable is included in equations (1) and (2).

3. We re-run equations (1) and (2) for shorter subsamples ending at T-23, T-22,..., then make 1, 4, 8 and 12-quarter ahead predictions, and compute and compare mean squared forecasting errors. Our forecast series contain, in general, 23, 20, 16 and 12 data points, respectively. However, in order to preserve enough degrees of freedom, the number of forecasts had to be reduced in those cases in which the indicator series does not cover the whole period.[3]

1.2 Financial indicators considered

In this paper, we analyse the informational content of 26 financial indicators, grouped in six different categories: domestic public debt yields, domestic public debt yield spreads, domestic-foreign interest rate differentials vis-à-vis Germany and the United States, credit quality spreads, exchange rates and stock prices. For comparative purposes, two standard monetary aggregates are also

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[2] See Appendix B for more details.

[3] See Appendix A for details regarding sample periods.

included: a narrow one (M2) and a broad one (ALP2). These financial indicators are fairly standard in the related literature.

As mentioned above, the intuition behind the use of financial indicators in this context is that forward-looking agents, when forming the expectations that determine financial prices, consider a wide information set. This information set includes not only the past course of fundamentals but also other pieces of information, such as monetary policy actions and their expected effects. It is precisely because of these *additional* pieces of information that financial indicators may have an *additional* information content compared to the macroeconomic fundamental above. The following paragraphs are not intended to provide a sound theoretical basis for the potential predictive power of each of the indicators considered. Such an analysis is beyond the scope of this paper.[4] Instead, these paragraphs are aimed at providing some insight into the potential predictive power of the chosen indicators.

In the first place, according to the Fisher relationship, domestic public debt yields can be decomposed into three unobservable components: the real interest rate, the expected rate of inflation over the life of the bond and the risk premia. To the extent that changes in yields reflect changes in the first component, they should be negatively correlated with future output growth. Similarly, changes in yields due to changes in the expected rate of inflation should, under reasonable assumptions, be positively correlated with future inflation.

The above-mentioned Fisher relationship can also explain why public debt yield spreads, defined as the difference between long and short yields, may contain significant information about future inflation. Regarding output, there are at least two possible explanations for the potential predictive power of the public debt yield spreads. The first is related to monetary policy. For example, a tightening of monetary policy, which will be followed by a fall in output growth, usually has a greater effect on short-term rates, flattening the yield curve. Alternatively, if agents are expecting low growth and they expect a Phillips curve relationship to hold, then inflation and interest rates would be expected to drop and the yield curve to flatten or even to invert. Notice also that, under the expectations hypothesis of the term structure of interest rates, yield spreads should be good predictors of future short yields.

Regarding the foreign-domestic interest rate differentials, if uncovered interest rate parity holds, these reflect the expected changes in the exchange rates. If purchasing power parity is also expected to hold, then expected exchange rate changes should be mirrored in expected inflation differentials. Thus, a wider differential may imply worse relative prospects for inflation in the home country. Moreover, both expected exchange rate changes and current exchange rates may have direct effects on output growth and, through this channel, on future inflation.

There are also two possible explanations for the potential predictive power of the credit quality spread, defined as the spread between the yield of a private asset and a public asset of the same maturity. First, since that spread should reflect mainly the greater default risk of the private asset, its changes could reflect changes in the perceived default risk, which should be negatively correlated with prospects of output growth. Second, Bernanke and other authors underline the relationship between the credit quality spread and monetary policy. According to these authors, in a context of imperfect sustitutability between assets, a monetary policy tightening induces a decline in the supply of bank loans. This means higher bank lending rates and higher rates on substitutes for bank loans, such as private bonds and commercial paper; i.e., a widening of the spreads between those rates and public debt yields. The predictive power regarding inflation could be based on a short-term relationship between output and inflation.

Finally, the use of stock prices can be justified as follows: since dividend growth will be related to output growth, stock prices can contain information about future output insofar as they reflect market expectations of future dividends.

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4 Woodford (1994), Davis and Fagan (1996), Estrella (1997) and Smets and Tsatsaronis (1997) provide a good basis for such a theoretical exercise.

2. Do financial indicators forecast inflation, output or short-term interest rates?

Regarding data, quarterly year-on-year CPI inflation, year-on-year GDP growth and 3-month domestic interest rates covering the period from 1978Q1 to 1997Q1 are the three macrofundamentals we consider. Details on the financial indicators considered are provided in Appendix A.

The main results of applying the process described in Section 1.1 to our data set are reported in Tables 1 to 3. Each table refers to one macrofundamental and shows which lags of the indicator are significant in the regression covering the whole period available, the number of observations in each equation, the ratio of the root of the mean in-sample squared error to that of the univariate model, and the mean squared error ratios corresponding to 1, 4, 8 and 12-quarters-ahead-out-of-sample forecasts. Two different values are provided for the last three ratios. First (upper values), ratios have been computed using the ex-post observed values of the indicator to make out-of-sample predictions. Second (lower values), out-of-sample values of the indicator have been forecast from an univariate equation containing 4 lags. The idea is that the actual predictive power of the indicator should be somewhere between the two ratios, because the univariate-based forecast of the indicator could be improved by a more general equation or model, but such an improvement would be limited by lack of perfect foresight.

Table 1 shows that only one term structure indicator is not significant in the equation for the inflation rate. According to the in-sample analysis, improvements vary between the 36% mean squared error reduction when the 5-year domestic yield (R5Y) is used and the 4% reduction corresponding to the 3-year domestic yield (R3Y). This result is similar to that found in most of the related papers for other countries. Out-of-sample results, however, are less favourable and, in general, ratios tend to be above 1. In 2 out of 8 cases the 1-quarter-ahead ratio is above 1. The best 1-quarter-ahead indicator is the 5-year yield (R5Y), which provides a ratio of 0.72. Results, however, are poorer for longer horizons. There are only three term structure indicators that offer ratios below 1 for four and eight quarters ahead projections and one regarding 12 quarters ahead. Only the 3-year to 1-month spread (S3_1) is able to outperform the univariate approach at any horizon, although the lowest ratio it provides is 0.89. Unfortunately, there are not sufficient data to test the out-of-sample performance of the more promising indicator according to the in-sample analysis: the 5-year domestic yield (R5Y).

Financial indicators based on the term structure offer by far the best results. Half of the domestic-foreign differentials are non-significant and those which are significant fail to improve the simple univariate results. Credit quality indicators tend to be significant but, when it is possible to make out-of-sample forecasts, these are outperformed by the univariate model. Similar results are obtained when using exchange rate and stock exchange indicators. It should be noted, however, that monetary aggregates do not provide better results, and have a poorer performance than the term structure indicators.

Overall, results in Table 1 raise some doubts about the usefulness of financial indicators as inflation predictors in Spain, at least for horizons between 1 and 12 quarters.[5] Are results similar regarding short-term interest rates and output?

According to Table 2, results are even worse regarding the 3-month interest rate. Although most indicators (18 out of 20) are significant in the regressions covering the whole period, their out-of-sample performance fails to provide ratios below 1. No indicator is able systematically to outperform the univariate model at any horizon. Only three indicators provide ratios below 1 for 1-quarter-ahead forecasts. This number falls to one for 4-quarter-ahead forecasts and to zero in the other two cases. Especially striking is the inability of long-term yields to provide good forecasts.

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[5] Slightly better results were obtained using an alternative price index (IPSEBENE by its Spanish name) which drops from the CPI the most volatile components.

Finally, Table 3 shows that many financial indicators are even non-significant in the regressions involving output (11 out of 26). Nevertheless, the 3-year domestic yield (R3Y) provides good results regarding the longest horizon and clearly outperforms the univariate model: the ratio for 12-quarter-ahead errors is 0.72 when the ex-post observed indicator is used and 0.68 when it is forecast with the univariate model. Similarly, the stock exchange indicator provides ratios below 1 for all horizons considered, varying between 0.75 and 0.95.

Table 1

The predictive power on inflation (CPI): linear model

Ind[1]	**Nobs.**	**Signif.[2]**	**Lags[3]**	**In-sample ratio[4]**	**Out-of-sample ratios[5]**			
					RMSE1	RMSE4	RMSE8	RMSE12
R1M	68	13.13	2-6[6]	0.93	0.99	1.00	0.99	1.10
		(0.04)				0.98	0.93	0.99
R12M	62	9.94	1-3[6]	0.95	1.00	1.07	1.05	1.10
		(0.04)				1.04	0.95	1.01
R3Y	63	5.46	-[6]	0.96	0.99	1.01	0.98	0.98
		(0.02)				1.06	0.98	1.03
R5Y	42	104.3	1-12[6]	0.64	0.72	–	–	–
		(0.00)				–	–	–
S5_1	43	41.12	6-12	0.70	1.28	1.21	–	–
		(0.00)				1.21	–	–
S3_1	61	35.64	6-11	0.86	0.89	0.89	0.95	0.94
		(0.00)				0.89	0.99	1.04
S12_1	61	22.15	2-10	0.89	0.98	0.93	1.00	1.03
		(0.01)				0.92	1.04	1.06
S5_12	43	21.41	9-12	0.80	0.89	1.04	1.22	–
		(0.00)				1.04	1.22	–
S3_12	63	6.47	1-5	–	–	–	–	–
		(0.26)				–	–	–
S12MG	59	27.10	1-12	0.91	1.08	1.16	1.03	0.97
		(0.01)				1.00	1.02	0.86
S3YG	60	14.04	1-12	–	–	–	–	–
		(0.30)				–	–	–
S5YG[7]	43	19.32	1-12	–	–	–	–	–
		(0.08)				–	–	–
S12MU	61	7.55	10-10	0.94	0.96	1.08	0.99	1.02
		(0.01)				1.08	0.99	1.06
S3YU	62	4.91	10-10	0.96	0.99	1.11	0.99	1.15
		(0.03)				1.11	0.99	1.17
S5YU	43	12.85	1-12	–	–	–	–	–
		(0.38)				–	–	–
SCP3M	31	55.59	3-8	0.68	–	–	–	–
		(0.00)				–	–	–
SCP12M	31	15.17	1-6	0.83	–	–	–	–
		(0.02)				–	–	–
SP5Y	45	4.05	3-6	–	–	–	–	–
		(0.40)				–	–	–
SCL3M	51	8.53	1-12	–	–	–	–	–
		(0.74)				–	–	–
SL3Y	64	13.48	6-8	0.93	1.17	1.10	1.18	1.02
		(0.00)				1.10	0.96	0.97
SL5Y	46	17.91	5-9	0.85	1.51	1.69	1.99	2.39
		(0.00)				1.69	1.91	1.10

Table 1 (cont.)

Ind[1]	Nobs.	Signif.[2]	Lags[3]	In-sample ratio[4]	Out-of-sample ratios[5]			
					RMSE1	RMSE4	RMSE8	RMSE12
ESPDEM	64	19.96	2-12	0.91	1.13	1.22	1.20	1.36
		(0.05)				1.27	0.98	0.96
ESPUSD	65	13.56	6-11	0.92	1.13	1.25	1.42	1.34
		(0.03)				1.25	1.28	0.89
NEER	64	21.75	2-12	0.88	1.23	1.51	1.77	1.97
		(0.03)				1.52	1.23	1.01
REER	64	21.25	1-12	0.88	1.39	1.84	2.09	2.18
		(0.05)				1.76	1.37	1.24
SP	66	18.19	3-10	0.89	1.15	1.19	1.25	1.28
		(0.02)				1.21	1.29	1.20
M2	65	16.62	5-11[6]	0.93	1.28	1.25	1.45	1.46
		(0.03)				1.24	1.38	1.32
ALP2	68	12.12	4-8[6]	0.94	1.02	1.20	1.22	1.09
		(0.06)				1.23	1.07	1.02

[1] See Appendix A for indicator definitions.

[2] Wald test robust to heteroscedasticity of the joint significance of the lagged terms of the indicator variable included in each equation. When cointegration exists, the null hypothesis also includes a zero value for the coefficient of the error correction term. The test has a $\chi^2(m)$ distribution, where m is the number of restrictions. p-value in parenthesis.

[3] Lagged terms of the indicator variable included in each equation.

[4] Ratio of one-quarter ahead RMSE, within sample, between the equation with indicator and the univariate equation. This ratio must always be smaller than one.

[5] Ratios of 1, 4, 8 and 12-quarters-ahead RMSE, out of sample, between the equation with indicator and the univariate equation. A value greater than one means worse forecast performance of the model with indicator than the univariate model. In general, in order to predict more than one quarter ahead, we need forecasts of the indicator itself. For each indicator, the first row is that resulting when actual values of the indicator are used for the forecasts and the second row is that resulting when AR(4) univariate predictions of the indicator are used. Results are presented only when at least 8 forecasts can be made.

[6] The model with indicator includes an error correction term, resulting from the cointegration between the levels of the dependent variable and the indicator.

[7] For this indicator, a trend is included in the equations, because only deviations of the indicator from a trend can be considered stationary.

All in all, the results in Tables 1 to 3 are rather negative regarding the ability of financial prices to forecast inflation, output or short-term interest rates. They seem to work, at least in most cases, when in-sample criteria are used but fail to do so out of the sample. This result is only partially at odds with other results in the literature which point to a higher informational content of financial indicators, because most of them are based solely on in-sample analysis.

Should we conclude that financial prices are not useful as indicators of future fundamentals in Spain? Before reaching such a conclusion, several aspects deserve more attention. Obviously, there are problems with the extension of some data series. But these problems can hardly be overcome unless we wait for about another ten years.

In our view, there are two more promising ways of gaining greater insight into the potential usefulness of financial prices. The first involves their usefulness as "qualitative" predictors. The idea is quite simple: maybe financial prices cannot anticipate the inflation rate prevailing, say, 2

Table 2

The predictive power on 3-month interest rates: linear model

Ind[1]	**Nobs.**	**Signif.**[2]	**Lags**[3]	**In-sample ratio**[4]	**Out-of-sample ratios**[5]			
					RMSE1	RMSE4	RMSE8	RMSE12
R12M	61	10.45	9-9[6]	0.92	1.08	1.33	1.42	1.54
		(0.01)				1.22	1.13	0.93
R3Y	59	27.74	7-12[6]	0.90	1.39	1.48	1.79	3.20
		(0.00)				1.41	1.60	1.63
R5Y	42	15.18	1-12[6]	–	–	–	–	–
		(0.30)				–	–	–
S12MG	61	25.31	1-10	0.83	1.37	1.51	1.64	2.31
		(0.00)				1.21	1.31	1.45
S3YG	60	10.22	9-12	0.93	1.04	1.14	1.20	1.91
		(0.04)				1.14	1.20	1.54
S5YG[7]	45	8.09	9-10	0.87	0.91	0.97	1.24	1.33
		(0.02)				0.97	1.24	1.21
S12MU	59	41.92	2-12	0.76	1.27	1.25	1.33	2.35
		(0.00)				1.30	1.23	1.26
S3YU	60	38.58	6-12	0.88	1.26	1.14	1.26	2.12
		(0.00)				1.14	1.37	1.79
S5YU	46	10.45	6-9	0.89	0.99	1.05	1.29	1.40
		(0.03)				1.05	1.42	1.14
SCP3M	34	19.15	1-5	0.90	1.17	–	–	–
		(0.00)				–	–	–
SCP12M	32	35.89	1-7	0.75	–	–	–	–
		(0.00)				–	–	–
SP5Y	42	16.89	4-11	0.87	0.97	1.96	–	–
		(0.03)				1.96	–	–
SCL3M	52	28.36	3-12	0.83	1.50	1.87	3.88	3.89
		(0.00)				1.48	1.61	1.29
SL3Y	63	11.90	6-9	0.93	1.06	1.12	1.15	1.31
		(0.02)				1.12	1.10	1.09
SL5Y	43	18.37	3-12	0.86	1.46	1.97	–	–
		(0.05)				2.10	–	–
ESPDEM	65	33.39	1-12	0.83	1.58	1.48	1.70	2.43
		(0.00)				1.33	1.24	1.28
ESPUSD	65	13.99	1-12	–	–	–	–	–
		(0.30)				–	–	–
NEER	67	17.19	6-10	0.92	1.17	1.06	1.15	1.46
		(0.00)				1.06	1.17	1.19
REER	69	18.32	1-8	0.90	1.44	1.48	1.69	2.27
		(0.02)				1.28	1.23	0.91
SP	68	19.73	1-9	0.95	1.23	1.19	1.22	1.61
		(0.02)				1.24	1.25	1.40
M2	72	12.52	2-2	0.94	1.00	0.91	0.61	0.94
		(0.00)				0.83	1.04	1.01
ALP2	65	16.13	1-12	–	–	–	–	–
		(0.19)				–	–	–

Note: For an explanation of the footnotes, see Table 1.

years ahead, but they can forecast whether prices are going to experience any unusual acceleration by that time. The second asks about the usefulness of financial prices as expectation indicators. We know that if expectations are rational and there are no information problems, expectations and ex-post values must differ only because of a standard white-noise term and, therefore, a good predictor will also be a good expectation indicator and vice versa. But in other perhaps more realistic circumstances,

Table 3

The predictive power on output: linear model

Ind[1]	**Nobs.**	**Signif.[2]**	**Lags[3]**	**In-sample ratio[4]**	**Out-of-sample ratios[5]**			
					RMSE1	RMSE4	RMSE8	RMSE12
R1M	67	22.64	4-10	0.87	1.08	1.04	1.06	0.82
		(0.00)				1.04	1.09	1.07
R12M	57	51.34	2-10	0.80	1.09	0.99	0.81	0.74
		(0.00)				1.01	0.91	1.03
R3Y	57	30.14	1-11	0.85	1.14	1.05	0.85	0.72
		(0.00)				1.07	0.90	0.68
R5Y	41	53.52	1-10	0.76	1.36	–	–	–
		(0.00)				–	–	–
S5_1	43	15.17	1-12	–	–	–	–	–
		(0.23)				–	–	–
S3_1	61	29.83	2-11	0.87	1.16	1.27	1.60	3.44
		(0.00)				1.26	1.52	3.85
S12_1	64	23.35	1-7	0.92	1.13	1.27	1.68	4.92
		(0.00)				1.26	1.56	5.17
S5_12	43	11.25	1-12	–	–	–	–	–
		(0.51)				–	–	–
S3_12	59	18.54	1-12	–	–	–	–	–
		(0.10)				–	–	–
S12MG	59	12.44	1-12	–	–	–	–	–
		(0.41)				–	–	–
S3YG	60	11.72	1-12	–	–	–	–	–
		(0.47)				–	–	–
S5YG[7]	43	57.00	4-12	0.66	1.12	1.31	–	–
		(0.00)				1.31	–	–
S12MU	65	5.47	2-2	0.96	1.07	1.14	1.47	4.49
		(0.02)				1.17	1.60	4.81
S3YU	63	31.37	2-9	0.85	1.08	1.08	1.25	2.14
		(0.00)				1.14	1.43	2.52
S5YU	45	33.14	6-10	0.83	0.91	0.93	0.86	–
		(0.00)				0.93	0.88	–
SCP3M	31	16.17	2-8	0.86	–	–	–	–
		(0.02)				–	–	–
SCP12M	31	7.83	1-8	–	–	–	–	–
		(0.45)				–	–	–
SP5Y	43	25.30	1-10	0.83	1.62	1.73	–	–
		(0.00)				1.66	–	–
SCL3M	52	21.98	1-12	0.89	1.17	1.25	1.39	–
		(0.04)				1.22	1.23	–
SL3Y	60	17.13	1-12	–	–	–	–	–
		(0.14)				–	–	–
SL5Y	43	16.13	1-12	–	–	–	–	–
		(0.19)				–	–	–

Table 3 (cont.)

Ind[1]	Nobs.	Signif.[2]	Lags[3]	In-sample ratio[4]	Out-of-sample ratios[5]			
					RMSE1	RMSE4	RMSE8	RMSE12
ESPDEM	65	25.65	3-12	0.87	1.09	1.14	1.36	2.57
		(0.00)				1.13	1.25	1.81
ESPUSD	65	14.94	1-12	–	–	–	–	–
		(0.24)				–	–	–
NEER	65	8.69	4-12	–	–	–	–	–
		(0.47)				–	–	–
REER	65	15.97	1-12	–	–	–	–	–
		(0.19)				–	–	–
SP	71	13.46	1-1	0.92	0.90	0.80	0.75	0.95
		(0.00)				0.86	0.91	0.92
M2R	65	20.54	3-12	0.90	1.35	1.22	1.33	3.55
		(0.02)				1.24	1.43	2.88
ALP2R	71	6.08	1-3	0.96	0.96	0.84	0.71	1.05
		(0.11)				0.88	0.85	0.96

Note: For an explanation of the footnotes, see Table 1.

even rational agents may be subject to important errors when predicting, for example, inflation. Therefore, indicators failing to forecast inflation might nevertheless be good inflation expectation indicators. In the next two sections we deal with these two issues.

3. Are financial prices useful as qualitative indicators?

In this section we explore whether financial prices are able to anticipate "events" although they are not able to ancitipate their "magnitude". If financial agents are forward-looking but tend to focus on general trends more than on eventual changes, financial prices would be better predictors of trend shifts than of precise point values.[6] This idea is behind the recent work by Estrella and Mishkin (1996) showing that the slope of the yield curve helps to predict recessions in the United States.

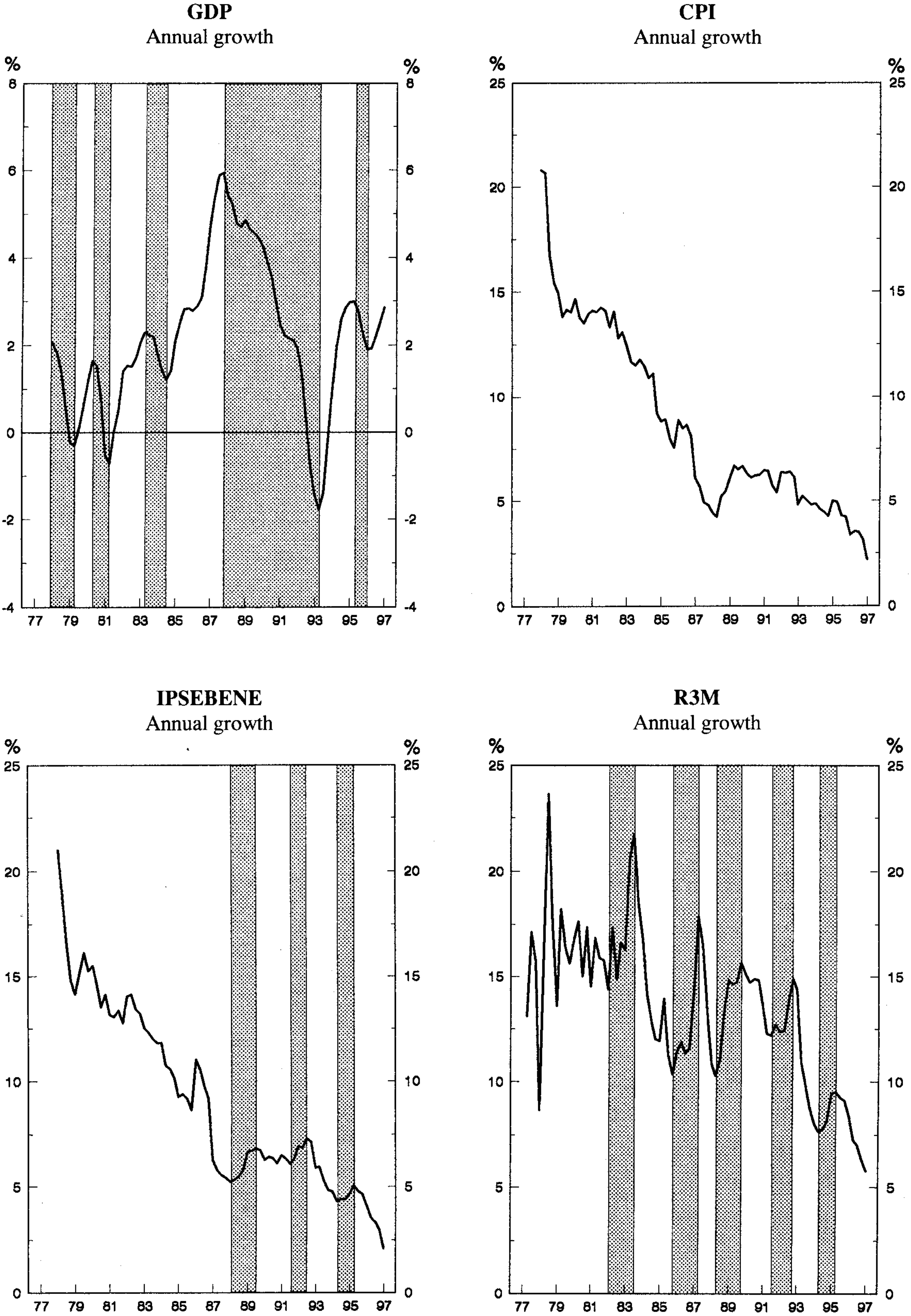
Exploring this possibility in detail is beyond the scope of this paper. Instead, we provide an initial approach for evaluating to what extent a deeper analysis might be worthwhile. Thus, we undertake a Probit analysis in which the qualitative dependent variables are "inflation upturns", "output slowdowns" and "monetary policy tightenings". Each of them has been built rather simply, following the procedure in Ball (1994). First, for inflation, output and the 3-month interest rate maxima (minima) are recorded as those observations that are higher (lower) than the three prior and the three subsequent observations.[7] Second, whenever two consecutive maxima (minima) are computed, the higher (lower) is chosen. Moreover, if there are two critical values separated by less than three quarters, the second one is eliminated. Finally, the dependent variables corresponding to

6 The fact that better quantitative results are obtained when a less volatile price index is used, see footnote 5, may be interpreted as providing some support for this view.

7 Regarding inflation, the less volatile index IPSEBENE has been used instead of CPI as an additional filter to eliminate noisy changes. Regarding output, the more classical approach of "three consecutive quarters of negative growth" has also been tried but it provided too few observations.

Chart 1

Macroeconomic variables

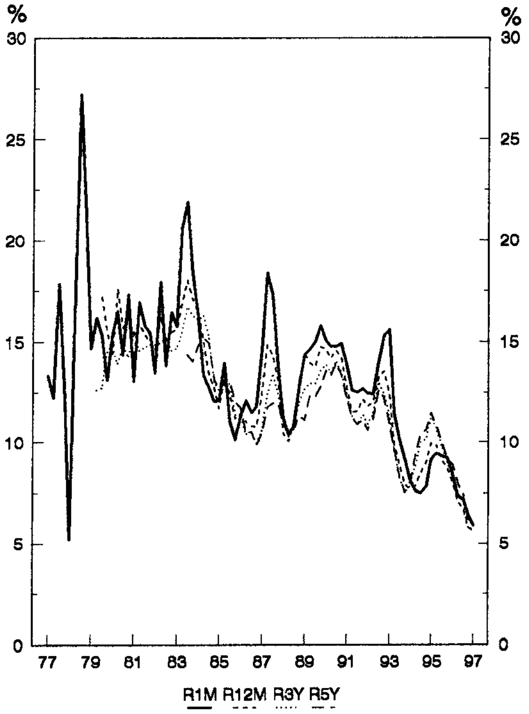


Note: Grey areas correspond to dates where the "qualitative" version of the variable is equal to one. See the main text for details.

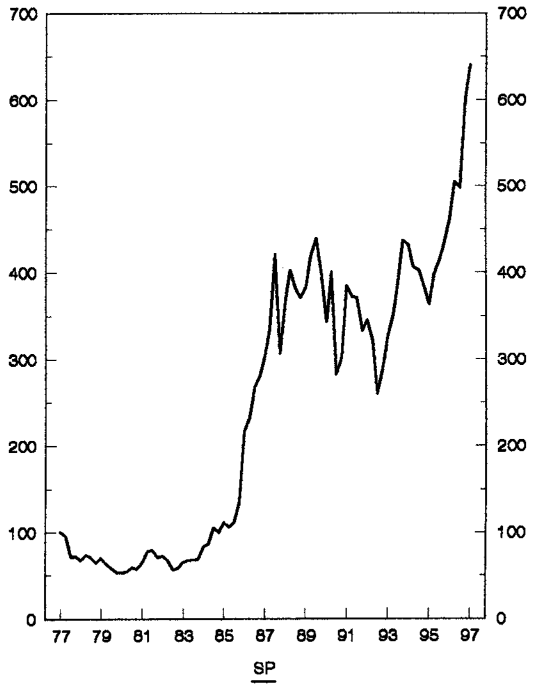
Chart 2

Indicators

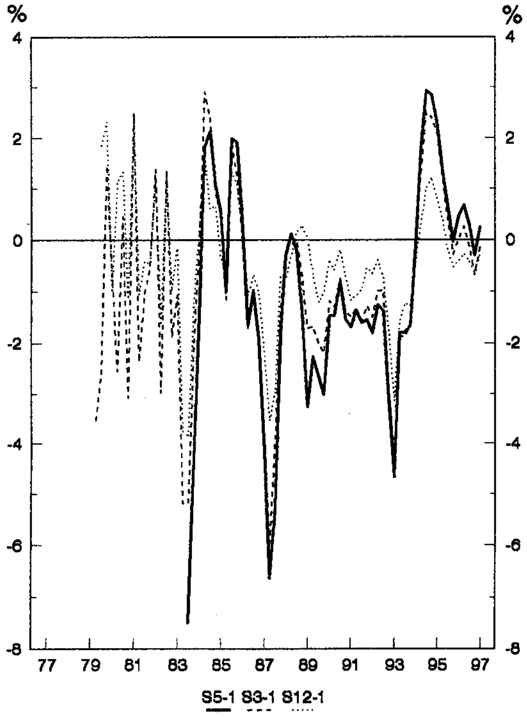
Riskless interest rates



Stock prices
Index (1977 1 = 100)



Term structure spreads
Against 1-month



Term structure spreads
Against 12-month

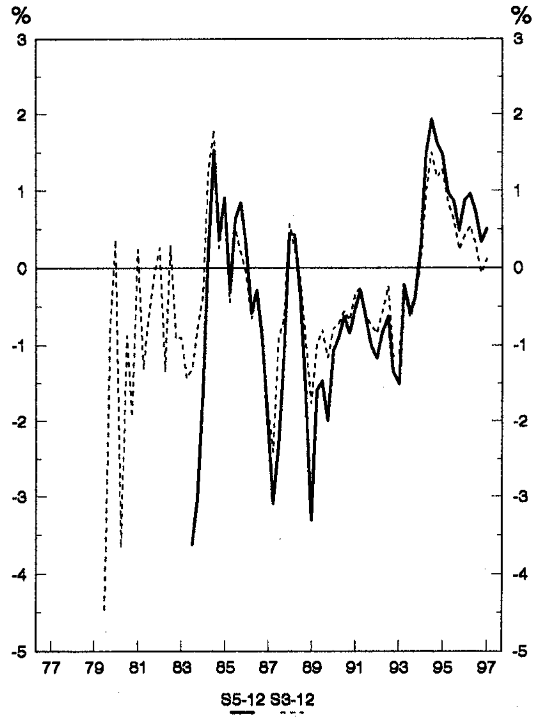
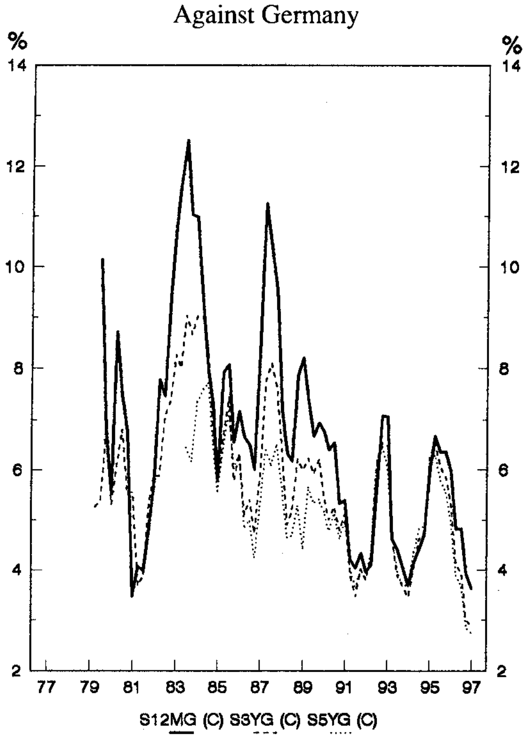


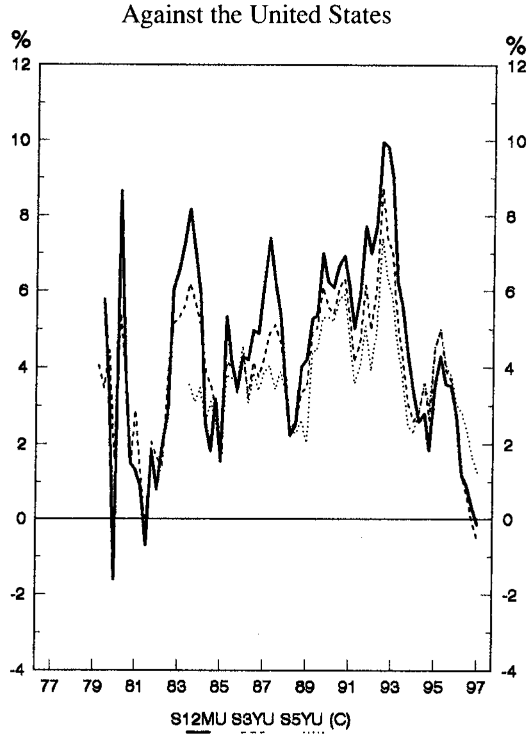
Chart 2 (cont.)

Indicators

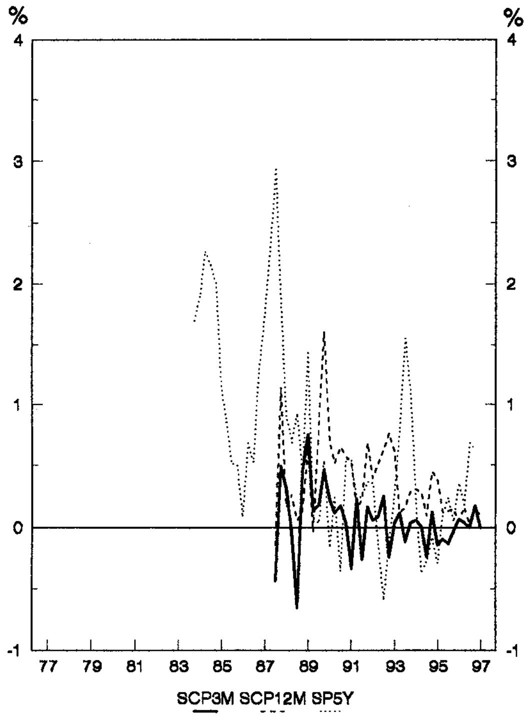
Foreign spreads
Against Germany



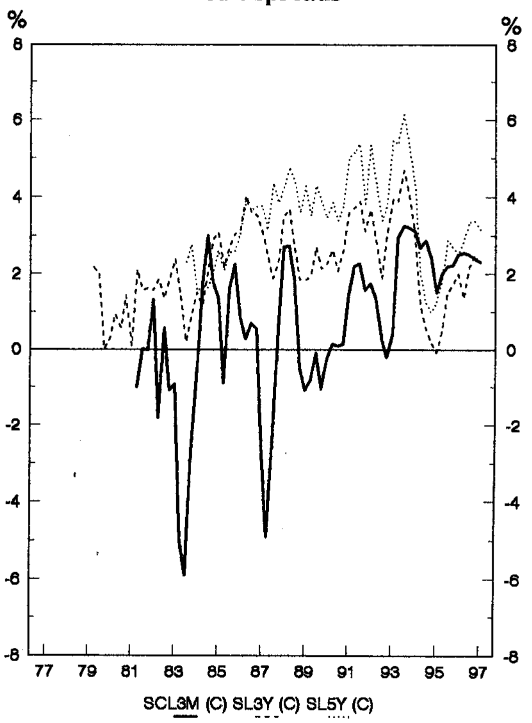
Foreign spreads
Against the United States



Private-public spreads



Credit spreads

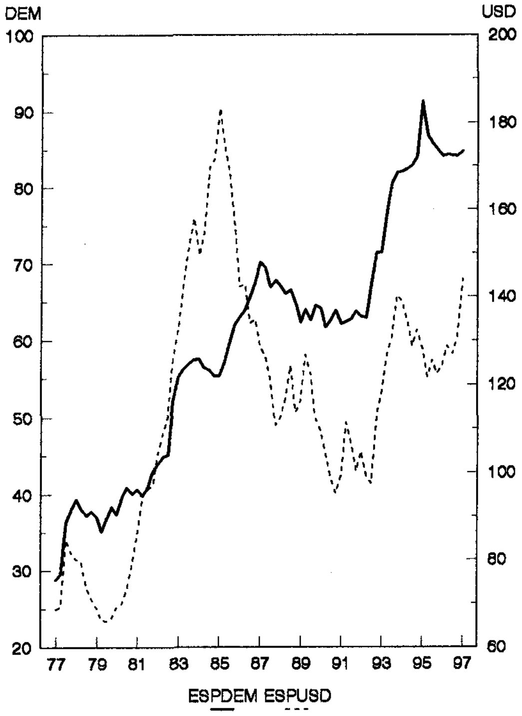


© Corrected series.

Chart 2 (cont.)

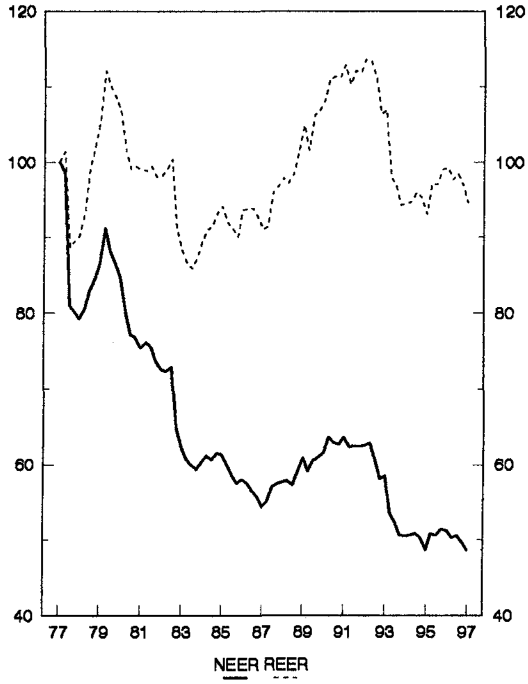
Indicators

Bilateral exchange rates



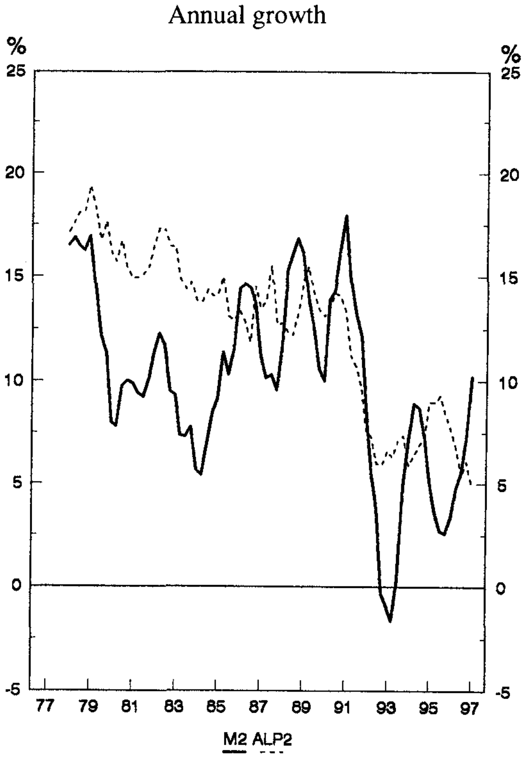
Exchange rate indices*

Index (1977 1 = 100)



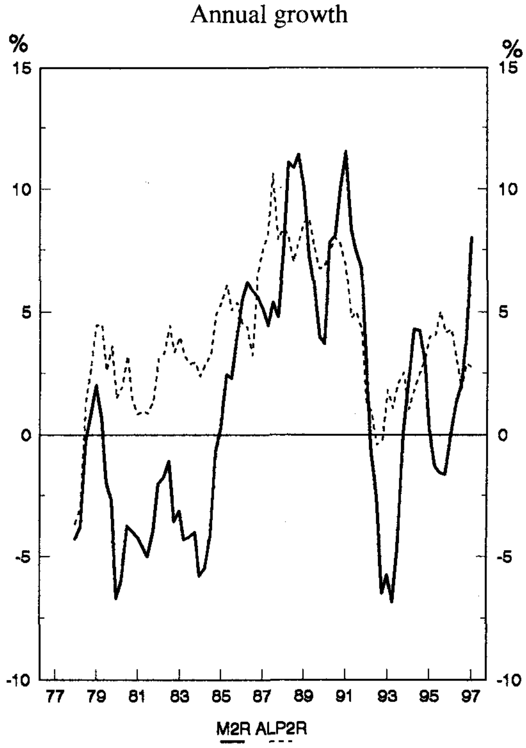
Nominal monetary aggregates

Annual growth



Real monetary aggregates

Annual growth



* A decline in the indices denotes a depreciation.

inflation and the interest rate are given the value of 1 whenever the corresponding series are moving from a minimum to a maximum. For output, values of 1 are given when it moves from a maximum to a minimum, thus reflecting a slowdown in output. Charts 1 and 2 show the variables used.

As to the Probit estimates and the performance criteria, they can be summarised in the following steps:

1. We first estimate a Probit model in which only (quantitative) lags of the fundamental are included. As before, this pseudo-univariate model will be our benchmark.
2. For those indicators that appeared as in-sample significant in the quantitative analysis, we add as many lags as suggested by the quantitative analysis.[8] The pseudo-R^2s suggested by Estrella (1995) and the mean probabilities corresponding to 1s and 0s are then compared. This is the equivalent of the in-sample quantitative analysis.
3. Both Probits are re-estimated for shorter samples and 23 1-quarter-ahead forecasts are made and compared according to the pseudo-R^2.

Tables 4 to 6 show the results of this procedure, which are rather promising. Regarding inflation, and in contrast to Table 1, most financial indicators that are significant in the in-sample analysis also have out-of-sample ratios below 1, what reflects a clear improvement over the univariate model. The higher increases in the pseudo-R^2 of out-of-sample forecast with respect to that of the univariate model correspond to the indicators based on the term structure: 3-year and 5-year yields (R3Y and R5Y) show ratios of 0.47 and 0.23, respectively; 5-year to 1-month (S5_1), 5-year to 1-year (S5_12) and 1-year to 1-month (S12_1) spreads also have low ratios (0.27, 0.40 and 0.55, respectively). Thus, financial indicators seem to do a better job forecasting inflation upturns than forecasting inflation itself.

Table 4

The predictive power on inflation (IPSEBENE): probit model

Ind[1]	**Nobs.**	**Signif.[2]**	**In-sample ratios**			**Out-sample ratio**
			P-R^2 [3]	Y=1 [4]	Y=0 [5]	P-R^2 [6]
R12M[7]	68	6.40 (0.01)	0.51	0.77	0.92	1.61
R3Y[7]	69	12.00 (0.00)	0.33	0.66	0.82	0.76
R5Y[7]	42	15.05 (0.00)	0.31	0.64	0.70	0.23
S5_1	45	13.88 (0.00)	0.30	0.67	0.72	0.27
S3_1	62	0.02 (0.90)	–	–	–	–
S12_1	46	5.33 (0.02)	0.60	0.81	0.92	0.55
S5_12	61	7.31 (0.01)	0.44	0.80	0.86	0.40
S3_12	59	0.01 (0.91)	–	–	–	–

8 In order to reduce the number of variables in the Probit model we consider a single variable built as an average of the different lagged values. Notice that the whole exercise is rather restrictive, which explains why this can be considered only as an initial approach.

Table 4 (cont.)

Ind[1]	**Nobs.**	**Signif.**[2]	**In-sample ratios**			**Out-sample ratio**
			P-R2 [3]	Y=1 [4]	Y=0 [5]	P-R2 [6]
S5YG	45	2.97 (0.08)	0.66	0.90	0.95	3.33
S3YU	68	0.04 (0.85)	–	–	–	–
S5YU	50	3.95 (0.05)	0.65	0.86	0.92	0.52
SCP3M	31	2.12 (0.15)	0.70	0.90	0.91	0.72
SCP12M	37	3.28 (0.07)	0.75	0.92	0.89	(-)
SP5Y	51	0.58 (0.44)	–	–	–	–
SL3Y	61	16.80 (0.00)	0.30	0.62	0.74	(-)
SL5Y	46	7.00 (0.01)	0.45	0.81	0.88	0.58
ESPDEM	73	0.84 (0.36)	–	–	–	–
ESPUSD	65	7.62 (0.01)	0.50	0.72	0.88	(-)
NEER	70	0.11 (0.75)	–	–	–	–
REER	64	1.26 (0.25)	0.85	0.96	0.99	0.73
SP	74	0.31 (0.58)	–	–	–	–

M2[7]	64	18.86 (0.00)	0.28	0.57	0.70	–
ALP2[7]	68	4.83 (0.09)	0.63	0.83	0.92	–

[1] See Appendix A for indicator definitions.

[2] Likelihood ratio test of the joint significance of the lagged terms of the indicator variable included in each equation plus the error correction term if this exists. The test has a χ^2 (m) distribution, where m is the number of restrictions. p-values in brackets.

[3] Ratio of pseudo-R^2, within sample, between the univariate equation and the equation with indicator. Within sample this ratio must always be lower than one.

[4] Ratio of the mean value of the fitted probability when Y is actually one in the univariate model and the model with indicator. A value lower than one implies that, on average, the model with indicator has a greater probability of being right when Y is equal to one.

[5] Ratio of the mean value of the fitted probability when Y is actually zero in the model with indicator and the univariate model. A value lower than one implies that, on average, the model with indicator has a greater probability of being right when Y is equal to zero.

[6] The same as footnote 3 for out-of-sample errors. The lower the ratio, the higher the informational content of the indicator. (-) denotes a negative ratio.

[7] The model with indicator includes an error correction term, resulting from the cointegration between the levels of the dependent variable and the indicator.

Table 5

The predictive power on output: probit model

Ind[1]	Nobs.	Signif.[2]	In-sample ratios			Out-sample ratio
			P-R2 [3]	Y=1 [4]	Y=0 [5]	P-R2 [6]
R1M	67	4.58 (0.03)	0.94	0.97	0.83	0.93
R12M	57	4.73 (0.03)	0.92	0.97	0.81	1.07
R3Y	57	11.33 (0.00)	0.85	0.93	0.58	4.35
R5Y	41	1.47 (0.23)	–	–	–	–
S3_1	61	9.37 (0.00)	0.88	0.93	0.72	0.90
S12_1	64	6.49 (0.01)	0.92	0.96	0.79	0.98
S5YG	43	0.23 (0.63)	–	–	–	–
S12MU	69	1.59 (0.20)	0.97	0.98	0.96	0.98
S3YU	63	0.45 (0.50)	–	–	–	–
S5YU	45	12.04 (0.00)	0.85	0.91	0.38	(-)
SP5Y	43	0.39 (0.53)	–	–	–	–
SCL3M	52	0.45 (0.50)	–	–	–	–
ESPDEM	65	1.51 (0.22)	–	–	–	–
NEER	65	2.28 (0.13)	0.97	0.98	0.94	0.93
SP	75	7.72 (0.01)	0.91	0.97	0.82	1.01
M2R	65	2.25 (0.13)	0.97	0.98	0.92	0.98
ALP2R	74	0.84 (0.36)	–	–	–	–

Note: For an explanation of the footnotes, see Table 4.

The same result applies to output slowdowns. According to Table 5, about half of the 9 significant indicators provide out-of-sample pseudo-R^2 ratios below 1. Again, the best results are provided by the yield slope indicators, the spread between 3 years and 1 month (S3_1) giving the lowest ratio: 0.90.

Similar results are found for the 3-month interest rate. In this case, 5 out of 9 significant indicators make better out-of-sample forecasts than the pure univariate model. It should be noticed again that the term structure appears as the more useful source of information. 1-year (R1Y) and 3-year (R3Y) yields are clearly able to outperform the univariate model, providing ratios of 0.74 and 0.64, respectively.

Table 6

The predictive power on 3-month interest rates: probit model

Ind[1]	Nobs.	Signif.[2]	In-sample ratios			Out-sample ratio
			P-R2 [3]	Y=1 [4]	Y=0 [5]	P-R2 [6]
R12M[7]	61	8.70 (0.01)	0.56	0.89	0.87	0.74
R3Y[7]	59	5.58 (0.06)	0.68	0.92	0.92	0.64
S12MG	61	0.53 (0.47)	–	–	–	–
S3YG	60	2.69 (0.10)	0.80	0.96	0.94	0.81
S5YG	45	6.45 (0.04)	0.58	0.89	0.83	5.26
S12MU	59	0.02 (0.90)	–	–	–	–
S3YU	60	1.29 (0.26)	0.89	0.97	0.98	0.94
S5YU	46	2.26 (0.13)	0.82	0.95	0.94	0.97
SCP3M	34	2.96 (0.09)	0.83	0.93	0.90	2.86
SCP12M	32	1.17 (0.28)	–	–	–	–
SP5Y	42	4.08 (0.04)	0.69	0.92	0.89	(-)
SCL3M	52	0.43 (0.51)	–	–	–	–
SL3Y	63	0.92 (0.34)	–	–	–	–
SL5Y	43	0.22 (0.64)	–	–	–	–
ESPDEM	65	0.10 (0.75)	–	–	–	–
NEER	67	0.03 (0.86)	–	–	–	–
REER	69	0.35 (0.56)	–	–	–	–
SP	68	4.01 (0.05)	0.68	0.93	0.92	3.33
M2	75	6.28 (0.01)	0.32	0.89	0.92	0.50

Note: For an explanation of the footnotes, see Table 4.

All in all, results in these last three tables are more promising than those of the quantitative analysis and point to the yield curve as a leading indicator of trend shifts in inflation, output and short-term interest rates.

4. May financial prices be useful as expectation indicators?

Up to now, we have focused on the capacity of financial prices to predict the future behaviour of some relevant macroeconomic variables. Nevertheless, even if they were bad predictors for those variables they might prove useful as indicators for agents' expectations. Obviously, it could be argued that rationality plus perfect information make this analysis redundant because expected and ex-post values only differ by a white-noise term. However, this rational-expectation-perfect-information framework is clearly at odds with what seems to be one of the main worries of most central bankers: the degree of credibility of the policies implemented. Naturally, credibility is a relevant issue only in a context of imperfect information.

A number of papers in the literature show how rational agents may be subject to important and rather persistent expectation errors. Most have focused on inflation. For example, some authors have found that, due to imperfect information, inflation rates can be successfully characterised by switching-regime models à la Hamilton, not only in high-inflation countries like Argentina, Israel or Mexico (see Kaminsky and Leiderman (1996)) but also in countries whose inflation rates are relatively low and stable like the United States (Evans and Lewis (1995)) or Canada (Bank of Canada (1996)). These switching-regime models produce inflation expectation errors which have zero mean *ex-ante* but, *ex-post*, can show a non-zero mean over relatively protracted periods. Similarly, according to King (1996), if agents do not immediately learn about central bank behaviour, disinflationary processes will probably be characterised by inflation targets (and, therefore, by actual inflation) below agents' inflation expectations. Lasting inflation expectation errors are also predicted by models à la Backus-Driffill (1985) where central bankers face credibility problems and need time to build their anti-inflationary reputation.

Differences between targeted values or planned monetary policy actions and expectations may imply additional costs to reach the targets or to implement the desired policy. For example, regarding inflation, discrepancies between targets and expectations, based on a credibility or information problem, may increase the costs of a disinflationary policy. Similarly, monetary authorities may provide clearer monetary policy signals if they know the interest rates agents are expecting. In these circumstances, agents' expectations are another valuable piece of information that financial prices could provide. In this section, we survey a number of recent papers on this issue written at the Research Department of the Banco de España.

The main problem in assessing the informational content of financial indicators in this respect is that agents' expectations are non-observable. Surveys, when available, rarely provide enough information. The approach, hence, has to be different. In particular, more room has to be made for economic theory and, arguably, results are model-dependent.

Our research in this area has been twofold. On the one hand, we have tried to retrieve inflation expectations from nominal interest rates according to the Fisher equation. On the other hand, expectations on future short-term interest rates have been obtained according to the relationship between short and long-term interest rates. As it is well known, however, an analysis of the informational content of financial prices on expected output cannot be based on similar non-arbitrage or equilibrium relationships.

The Fisher relationship states that riskless nominal interest rates are equal to the sum of three components: a riskless real rate to the same maturity, the expected inflation at that horizon and an inflation risk premium. If we do not believe there are arbitrage opportunities in Spanish financial markets, inflation expectations at different horizons could be obtained provided we have data on the nominal zero-coupon bond yield curve, the real zero-coupon bond yield curve and the inflation risk premia for different maturities.

The nominal zero-coupon yield curve is regularly estimated at the Banco de España following the Nelson and Siegel (1987) and Svensson (1994) methodology. This method provides a smooth continuous nominal zero-coupon yield curve, and according to Núñez (1995) offers better results than alternative methods available in the literature.

In Ayuso (1996), *ex-ante* real rates are estimated for the Spanish economy in a CCAPM framework. Notice that *ex-post* real interest rates are not good substitutes for *ex-ante* interest rates in this case for, at least, two reasons. For one thing, the Fisher relationship would imply that the average inflation risk premium is zero. For another, *ex-post* real interest rates are only observable after the inflation rate has been observed, thus dispelling any usefulness they may have as an indicator of inflation expectations. Therefore, *ex-ante* real interest rates have to be estimated.

The approach in Ayuso (1996) can be briefly summarised as follows. For the equilibrium relationships implied by the CCAPM for returns expressed in real terms, it can be shown that the riskless zero-coupon *ex-ante* real interest rate to a given horizon k must be equal to the inverse of the expected marginal rate of substitution between current and k-period-ahead consumption. If agents have isoelastic preferences and consumption and returns are jointly lognormal, the marginal rate of substitution depends on two parameters that characterise agents' time preference and risk aversion, respectively, and the (log) rate of consumption growth.

The time preference and the relative risk aversion parameters are estimated following Hansen and Singleton (1982): without imposing lognormality, first-order conditions for different investment strategies maturing between 1 and 12 months in the future are obtained. In particular, for each maturity, several combinations of 1 to 12-month zero-coupon bonds are considered. This set of first-order conditions is then used to estimate, by GMM, the above-mentioned parameters. Expected consumption growth at different horizons are obtained from an AR-ARCH model for consumption growth. Table 7 shows the basic statistics thus obtained for the 1, 3, 5 and 10-year *ex-ante* real interest rates. As can be seen, they seem to be rather stable and the real yield curve is nearly flat. It should be said, however, that the level of the real yield curve is not estimated with high precision.

Table 7

Basic statistics of *ex-ante* real interest rates

Maturity	Minimum	Maximum	Average	Standard deviation
1 year	3.92	5.67	4.88	0.28
3 years	4.42	5.21	4.85	0.13
5 years	4.57	5.06	4.84	0.08
10 years	4.70	4.94	4.83	0.04

Notes: Taken from Ayuso (1996). Rates measured in annual percentage points (log approximations). Data are monthly and cover the period 1985:2-1994:12.

Turning now to inflation risk premia, an estimate is undertaken in Alonso and Ayuso (1996) also in a CCAPM framework assuming both lognormality and isoelastic preferences. Under these assumptions, it is easy to show that for any horizon k the inflation premium can be expressed as the product of two factors: the agents' relative risk aversion coefficient, and the conditional covariance between k-period-ahead (log) prices and consumption. They estimate 1, 3 and 5-year-ahead conditional covariances between Spanish price and consumption data from a bivariate GARCH model and calculate inflation premia for different available estimates of the Spanish relative risk aversion coefficient. Table 8 shows the basic statistics for the inflation premia when the maximum estimate of relative risk aversion (7.22) is considered. This can be seen as an upper bound for the actual inflation premia. According to this table, inflation premia can also be considered relatively low and stable even for maturities up to 5 years.

Regarding the informational content of the term structure for inflation expectations, the above results suggest that, since the level of the real yield curve is estimated with low precision, the most efficient way to exploit the informational content of long-term nominal interest rates is by looking at changes in their levels. Given that inflation premia and *ex-ante* real rates are rather stable, changes in long-term zero-coupon interest rates should mainly reflect changes in agents' inflation expectations.

Table 8

Basic statistics of inflation premia

Inflation premium at	Minimum	Maximum	Average	Standard deviation
1 year	0.02	0.04	0.03	0.004
3 years	0.09	0.20	0.11	0.029
5 years	0.18	0.39	0.23	0.052

Notes: Taken from Alonso and Ayuso (1996). In annual percentage points (log approximations). Data are quarterly and cover the period 1973:I-1995:IV.

As to the possibility of extracting information for short-term interest rate expectations from long-term interest rates, it is well known that long-term rates can be expressed as an average of future expected short-term rates plus a term risk premium. Term premia for equations containing expectations on 1-month and 1-year interest rates to different horizons have been estimated in Restoy (1995), using the methodology proposed by Backus and Zin (1994) to explain the shape of a yield curve.

The starting point of this methodology is a non-arbitrage argument: if there are no arbitrage opportunities, all expected returns must be equal provided they are discounted using the proper discount factor. Assuming that the discount factor follows an ARMA process, it is easily shown that the parameters of this process completely characterise the current interest rates, the implicit forward rates and the term premia. Thus, term premia can be computed, provided estimates of the ARMA parameters are available. The discount factor, however, is non-observable and this precludes the direct estimation of its univariate model. But the ARMA parameters can be retrieved, exploiting the fact that they also determine the sample moments of current and forward interest rates.

This retrieval process is what Backus and Zin (1994) call a "reverse engineering process": given an autoregressive order and a moving average order, the relationship between the ARMA parameters of the process followed by the discount factor and the sample moments in the time series of the spot and forward interest rates can be used to estimate the former from the latter. Different AR and MA orders give rise to a different set of parameters and GMM provides a natural way of, first, estimating them, and second, choosing the model that best fits the data.

Table 9

(Average) term premia

Within	Term premium corresponding to		Pro memoria: average	
	1-month interest rate	1-year interest rate	1-month forward rate	1-year forward rate
1 month	0.01	0.01	10.63	10.49
3 months	0.03	0.03	10.89	10.47
1 year	0.12	0.11	10.45	10.48
3 years	0.38	0.30	10.60	10.60
5 years	0.47	0.44	10.56	10.53
10 years	0.70	0.55	10.22	10.20

Notes: Taken from Restoy (1995). Annualised premia and rates, in percentage points (log approximations). Data are monthly and cover the period 1991:1-1995:7.

Table 9 presents the average term premium estimates obtained in Restoy (1995), together with the mean values of the 1-month and the 1-year forward rates. According to the estimates in the table, term premia included in Spanish nominal interest rates can be considered moderate or low, and therefore, 1-month and 1-year forward curves, which are obtained from the zero-coupon nominal yield curve- can be seen as mainly reflecting the expected paths for 1-month and 1-year interest rates.

Conclusions and policy implications

In this paper we have analysed the informational content of different financial prices on three macroeconomic variables of clear interest to the Banco de España in the design and implementation of its monetary strategy: the inflation rate, i.e. the direct target of current Spanish monetary policy; a short-term interest rate, i.e., its operational target; and output, because even a central bank with direct final inflation targets should worry about output deviations from a reference level.

We have looked at 26 financial prices covering the term structure, foreign-domestic differentials, credit quality, exchange rates and stock exchange indicators and have checked, first, their capacity to forecast quantitatively the three above-mentioned macrofundamentals; second, their usefulness as "qualitative" predictors of inflation upturns, output slowdowns and monetary policy tightenings; and, finally, their usefulness as inflation and interest rate expectation indicators. In some sense, and guided by the results, we have moved from a very demanding to a less demanding analysis.

Although most of the financial indicators considered are found to be significant when included in the regression to explain the behaviour of inflation, output or the interest rate, they fail to outperform a simple univariate model when their out-of-sample performance up to three years is analysed.

Given this result, we have explored the possibility of using those financial indicators as "qualitative" rather than as "quantitative" indicators. As an initial approach, we have estimated several Probit models to forecast inflation upturns, output slowdowns and monetary policy tightenings. The results of this approach are clearly promising and seem to merit a further analysis that is beyond the scope of this paper. In any case, they point to the yield curve as the main potentially useful source of information.

Finally, we have also explored whether financial prices may be considered as good expectation indicators, irrespective of their ability as quantitative or qualitative predictors. The rationale for this analysis is based on agents' inability to perceive clearly what central banks really do. In this framework, they could make errors that are far from the usual zero-mean assumption. Although the approach relies on the acceptance of several prior assumptions, the available evidence points to an important informational content of yields on zero-coupon bonds on both expected inflation and expected short-term interest rates.

Taken together, these results may have important implications for the use of financial indicators in the current Spanish monetary policy framework. As none of the financial indicators considered seems to hold a stable empirical relationship with any of the fundamentals, this discards the possibility of using them as nominal anchors for monetary policy decisions in the same way that monetary aggregates were used in the past. Nevertheless, they can be useful both as "qualitative" indicators to complement the quantitative information provided by other non-financial indicators, and as expectation indicators signalling potential credibility problems and potential misunderstandings of monetary policy actions. In this respect, indicators derived from the zero-coupon yield curve (interest rate levels and spreads) emerge as the most informative financial prices.

Appendix A: Data description

Due to the late development of a full range of liquid and competitive financial markets, the availability of data on asset prices in the Spanish economy is very limited. As a consequence, the selection and construction of variables for this work has been influenced by the need to have information for a period long enough to make reliable estimations of information content. This means that, in some cases, the variables used are only an approximation to the theoretical variable of interest.

In this appendix we describe the variables used in this work.[9] Unless otherwise indicated the source is the Banco de España and the quarterly series are built as the monthly averages of the daily data corresponding to the last month of each quarter. Most series cover the period from the first quarter of 1977 to the first quarter of 1997, but some of them do not cover the whole period.

Macroeconomic variables:

GDP: Real Gross Domestic Product. Source: National Institute of Statistics (INE). Quarterly series in origin.

CPI: Consumer Price Index. This is a re-elaboration, made at the Banco de España, of the index produced by the National Institute of Statistics (INE) to homogenise the methodology of calculation for the whole period. Monthly in origin.

IPSEBENE: Consumer Price Index corrected by the elimination of its more volatile components: energy and non-processed foods. As before, we use the series re-elaborated at the Banco de España. Monthly in origin.

R3M: 3-month interbank interest rate.

Domestic riskless interest rates:

R1M: 1-month interbank interest rate.

R12M: 12-month interbank interest rate.

R3Y: 3-year central government bond yield. Until 1988, average yield on outright spot transactions with bonds at between 2 and 4 years on the Madrid Stock Exchange. Thereafter, average yield on outright spot transactions between market members with 3-year bonds on the public debt Book-Entry Market.

R5Y: 5-year central government bond yield. Until 1991, average yield on bonds at over 4 years. Thereafter, average yield on 5-year bonds. Data from outright spot transactions between market members on the public debt Book-Entry Market since 1988 and from the Madrid Stock Exchange before then.

Term structure spreads:

S5_1: 5-year minus 1-month (R5Y-R1M).

S3_1: 3-year minus 1-month (R3Y-R1M).

S12_1: 12-month minus 1-month (R12M-R1M).

S5_12: 5-year minus 1-year (R5Y-R12M).

S3_12: 3-year minus 1-year (R3Y-R12M).

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9 All of them are shown in Charts 1 and 2.

Domestic-foreign spreads:

S12MG: 12-month interbank interest rate in Spain (R12M) minus 12-month interbank interest rate in Germany. Domestic markets.

S3YG: 3-year government bond yield in Spain (R3Y) minus 3-year government bond yield in Germany.

S5YG: 5-year government bond yield in Spain (R5Y) minus 5-year government bond yield in Germany.

S12MU: 12-month interbank interest rate in Spain (R12M) minus 12-month interbank interest rate in the United States. Domestic markets.

S3YU: 3-year government bond yield in Spain (R3Y) minus 3-year government bond yield in the United States.

S5YU: 5-year government bond yield in Spain (R5Y) minus 5-year government bond yield in the United States.

Credit quality spreads:

a) Private-public spreads:

SCP3M: 3-month commercial paper interest rate minus 3-month Treasury bill interest rate. In both cases, interest rates correspond to primary auction markets. Only auctions of the major issuers are considered. These are semi-public companies, but they are the only ones that conduct auctions regularly.

SCP12M: 12-month commercial paper interest rate minus 12-month Treasury bill interest rate. Comments on the previous variable also apply here.

SP5Y: Corporate bond yield minus 5-year government bond yield. Average yields in secondary markets. Corporate bonds correspond to electric companies and have horizons of about 2 years.

b) Credit spreads:

SCL3M: Average interest rate of banks and savings banks on commercial discount up to 3 months minus 3-month interbank interest rate (R3M).

SL3Y: Average interest rate of banks and savings banks on credit accounts at 1 to 3 years minus 3-year government bond yield (R3Y).

SL5Y: Average interest rate of banks and savings banks on loans at 3 years or over minus 5-year government bond yield (R5Y).

Exchange rates:

ESPDEM: Spot price of the Deutsche mark in pesetas per unit.

ESPUSD: Spot price of the US dollar in pesetas per unit.

NEER: Index of the nominal effective exchange rate of the peseta against developed countries.

REER: Index of the real effective exchange rate of the peseta against developed countries.

Stock prices:

SP: Madrid Stock Exchange General Index, end-of-month data. Source: Madrid Stock Exchange.

Monetary aggregates:

M2: Narrow measure of money in nominal terms.

ALP2: Broad measure of money in nominal terms. The original series is adjusted for a change in level at the beginning of 1992, due to the exchange of Treasury notes for especial public debt.

M2R: M2 deflated by CPI.

ALP2R: ALP2 deflated by CPI.

Appendix B: Unit root test and data transformations

We make several transformations of the original data. First, all interest rates, and consequently all spreads, are expressed in continous time. Second, the rest of the series are expressed in logarithms. Finally, all series are duly transformed to include only stationary series in the equations. This last step requires the analysis of the order of integration of the different variables considered, as well as the possible existence of cointegration relationships between some of them.

Most variables considered have been frequently used in empirical work. Thus, there is widespread evidence about their univariate and bivariate stochastic properties. Consequently, we shall not repeat here the analysis of those variables, but concentrate on those less frequently analysed.

Summarising previous evidence, we know that both price indices (CPI and IPSEBENE) are seasonal I(2) variables, so a $\Delta\ \Delta_4$ transformation in logarithms ensures stationarity (see, for example, Matea and Regil (1996)). GDP is a borderline case between I(1) and I(2), depending on the particular sample period considered. In this work, we considered GDP as I(1). Although, by construction, GDP should be a nonseasonal variable, there is some evidence of seasonality in it. So, we use a Δ_4 of the log of GDP as the stationary transformation.

As regards interbank and public debt interest rates, Alonso et al. (1997) have shown that they are I(1) variables, that they are cointegrated with the annual growth of both price indices and that spreads between them are stationary.

Likewise, the different exchange rates considered are I(1) variables. This result also applies to the real effective exchange rate index, which implies the non-existence of cointegration between the nominal effective exchange rate and consumer prices (see Pérez-Jurado and Vega (1993)).

Finally, nominal monetary aggregates are I(2) but real monetary aggregates are I(1) and all of them have seasonal components. That is, the growth rate of nominal monetary aggregates and inflation are cointegrated (see, for example, Ayuso and Vega (1994)).

Regarding the remaining indicators considered in this work (domestic-foreign, private-public and credit spreads), we present here some evidence about their stochastic properties. Initial tests showed the existence of a unit root in some of these spreads. But the low power of these test against the alternative of stationarity with some structural break is well known. In fact, the Spanish economy, and its financial system in particular, has experienced significant changes over the sample period considered.

A quick look at the series suggests specific dates at which a change in the mean occurs for several related series. Hence, we observe a change in the mean of the credit spreads around 1984:4, probably reflecting the passing from a context of legally fixed banking rates to one of market-determined rates.[10] Similarly, the recent convergence of Spanish interest rates towards the German

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[10] The liberalisation of interest rates on bank assets began in 1977 and was completed in 1981. Interest rates on bank liabilities were not fully liberalised until 1987.

ones can be represented as a change in the mean of Spanish-German spreads around 1991:1. We eliminate these changes in the mean from the original series, using univariate models to estimate the corrected series. More statistically than theoretically grounded is the correction in the spread between Spanish and US 5-year rates for a change in the mean in 1996:2.

Table B.1

Unit root tests: I(1) against I(0)

Model with trend				
	S12MG(c)	S3YG(c)	S5YG(c)	S12MU
$t_{\tilde{\alpha}}$	-3.30*	-3.21*	-3.06	-2.48
ϕ_3	5.54*	5.51	4.87	6.36*
ϕ_2	3.80	3.71	3.41	4.25*
	S3YU	S5YU(c)	SCP3M	SCP12M
$t_{\tilde{\alpha}}$	-2.00	-2.54	-7.15***	-5.67***
ϕ_3	4.06	3.54	26.49***	16.80***
ϕ_2	2.74	2.40	17.69***	11.21***
	SP5Y	SCL3M(c)	SL3Y(c)	SL5Y(c)
$t_{\tilde{\alpha}}$	-2.88	-3.71**	-3.20*	-2.46
ϕ_3	4.75	6.98**	5.21	3.20
ϕ_2	3.22	4.69*	3.47	2.14
Model without trend				
	S12MG(c)	S3YG(c)	S5YG(c)	S12MU
$t_{\alpha*}$	-2.97**	-2.54	-2.02	-3.74***
ϕ_1	4.60*	3.31	2.33	7.17***
	S3YU	S5YU(c)	SCP3M	SCP12M
$t_{\alpha*}$	-2.95**	-2.67*	-6.35***	-5.10***
ϕ_1	4.51*	3.68	20.74***	13.34***
	SP5Y	SCL3M(c)	SL3Y(c)	SL5Y(c)
$t_{\alpha*}$	-2.66*	-3.11**	-3.03**	-2.54
ϕ_1	3.68	4.97**	4.64*	3.29

Notes:
1. A (c) indicates that the corrected series has been used.
2. *, ** and *** indicates significance at the 10%, 5% and 1% levels, respectively.
3. Both models contain a constant and 4 lags of the corresponding spread.
4. t_α is a test of the mull hypothesis of existence of a unit root in the corresponding model.

Table B.1 shows Phillips-Perron unit root tests[11] for foreign, private-public and credit spreads. When needed, the corrected serie is used. With a few exceptions, the existence of a unit root can be rejected for all series, at least at the 10% significance level. When not significant, the statistics are very close to the 10% critical value (in the model with trend for the case of the 5-year spread with Germany).

[11] For details about the calculation and interpretation of the tests, see Perron (1988).

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Interpretation of the information content of the term structure of interest rates

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Introduction

The objective of this paper is to analyse the information content of the term structure of interest rates in Belgium. It is, however, well known that the intermediate target of Belgian monetary policy is the stabilisation of the DM/BF exchange rate. This type of policy has produced close links between Belgian and German interest rates and therefore between both countries' term structures. Hence, the analysis of the Belgian term structure cannot be isolated from what happens in Germany and our analysis is therefore extended to include the information content of the term structure of German interest rates as well.

1. Responses of market rates to key central bank interest rates

A popular approach to the term structure of interest rates is the expectations hypothesis. According to this theory and assuming a constant risk or liquidity premium, a longer term interest rate (in terms of spot rates or spot yields) can be written as an average of an actual shorter term spot rate with given maturity and expected future values of that same short rate up to the maturity of the longer rate. The shortest available interest rate is the overnight rate, which is closely linked to the official central bank interest rate. Therefore any longer term spot rate can be written as an average of the actual and future expected values of the overnight rate. Consequently, any interest rate can, according to the expectations hypothesis, be considered to reflect market participants' expectations concerning the future stance of monetary policy. The reaction of market rates to changes in official rates may thus reveal the degree to which the market correctly anticipates future policy moves.

The reaction of market rates to changes in central bank rates may be less than proportional and decline with the maturity of the asset. Equivalently, the implicit forward rates further in the future may react less or even inversely to the change in the official interest rates. This result can be explained by two arguments: first, interest rates, and especially the official target rates, follow a mean-reverting process, so that long rates should react less then short rates. Changes in the official rates are persistent but not permanent; second, some part of the change in the official rate may already be anticipated by the market, such that even short market rates may not fully adjust to official rate changes.

It is therefore interesting to decompose the reaction of market interest rates into the anticipated component and the surprise or announcement effect, on the one hand, and to detect the degree of persistence in market expectations on the other. This decomposition can reveal how correctly the market understands the reaction function of the central bank, and whether it considers the official interest changes as permanent or as temporary. Perhaps this last issue can be interpreted as a measure of credibility. If the policy is thought to be effective, official interest rate changes (to bring a given objective of monetary policy such as inflation or, as in Belgium, the DM exchange rate, back to its target value) will be considered to be temporary in nature. If, on the other hand, the market participants consider the policy move to be ineffective or incredible, it will obtain a permanent or even extrapolative character.

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* The opinions expressed in this paper do not necessarily represent those of the National Bank of Belgium.

Following Roley and Sellon (1996), we consider the relation between the target value of the monetary policy objective in the following month (X_{t+1}) to be related to the actual one-month interest rate ($r_{1,t}$) and a random error term (u_{t+1}):

$$X_{t+1} = a - br_{1,t} + u_{t+1} \tag{1}$$

The random error is autocorrelated:

$$u_{t+1} = \rho u_t + \nu_{t+1} \tag{2}$$

where ν is a white noise process and ρ measures the degree of persistence of economic or financial shocks.

Market participants expect the central bank to influence the money market rate in order to hit its target of monetary policy. They therefore expect, for example, the one-month interbank rate to reflect the central bank's reaction function as:

$$r_{1,t}^{a} = \alpha + \beta\left[E(X_{t+1}) - \overline{X}\right] + \varepsilon_t \tag{3}$$

where $r_{1,t}^{a}$ is the one-month interest rate just after the policy move at the end of period t (in the empirical work this will be equivalent to the market rate the day after the official rate has changed).

From the preceding equations, the market participants' perceived move (abstracting from uncertainty) of the one-month interest rate between the end of period t-1 and the end of period t is:

$$r_{1,t}^{a} - r_{1,t-1}^{a} = \delta\left[\varepsilon_t - \varepsilon_{t-1} + \beta\rho(\rho u_{t-1} + \nu_t - u_{t-1})\right] = \Delta r_t^{o} \tag{4}$$

where $\delta = \dfrac{1}{(1+\beta b)} < 1$ and r^{o} is the official interest rate.

However, there is always a probability that the central bank does not follow its reaction function and undertakes a discretionary action. If market participants are sure that the central bank will follow its reaction function and therefore change its official rate in line with deviations between actual and target values of the central bank's objective, the market rate will anticipate such a rise before the official move:

$$r_{1,t}^{b} - r_{1,t-k}^{a} = (1-\theta)\Delta r_t^{o} \tag{5}$$

where r_1^{b} is the one-month rate just before the official rate move (in the empirical work the day before an official rate change) and θ is the probability that the central bank will not change the official rate (i.e. will act in a discretionary way).

The immediate response of the one-month market rate to the policy change (i.e. the difference between the market rate at the day after and the day before the policy move) is then:

$$r_{1,t}^{a} - r_{1,t}^{b} = \theta\Delta r_t^{o} \tag{6}$$

For longer maturity market rates the relation becomes more complicated:

$$r^a_{N,t} - r^b_{N,t} = (\frac{1}{N})\sum_{j=0}^{N-1}\theta^{j+1} * \Delta r^o_r \tag{7}$$

$$r^b_{N,t} - r^a_{N,t-k} = (\frac{1}{N}) * \left[(1-\theta) + \sum_{j=1}^{N-1}(\rho^j - \theta^{j+1} + (1-\rho)\sum_{i=1}^{j}\theta^i \rho^{j-i}) \right] * \Delta r^o_t \tag{8}$$

In the case of N=3, this becomes:

$$r^a_{3,t} - r^b_{3,t} = (\frac{1}{3}) * (\theta + \theta^2 + \theta^3) * \Delta r^o_t$$

$$r^b_{3,t} - r^a_{3,t-k} = (\frac{1}{3}) * \left[(1-\theta) + (\rho - \theta^2 + (1-\rho) * \theta\ ) + (\rho^2 - \theta^3 + (1-\rho) * \theta\ \ \rho^1 + (1-\rho) * \theta^2) \right] * \Delta r^o_t$$

Estimation over the period February 1991 (start of the new organisation of the monetary policy in Belgium) to August 1996 (the penultimate change in the official rate) results in a non-significant negative coefficient for ρ. However leaving out the period of 23rd July –15th September 1993 where the official rate was raised to defend the Belgian franc against speculative pressures, the estimates produce a significant and high value for ρ = 0.8. The changes of the central rate are thus considered by the market as persistent but non-permanent movements.

The estimation of the θ coefficient, measuring the degree of surprise (or, inversely, the degree of anticipation) of interest changes, gives far fewer problems. In all cases it is situated around 0.84 and is highly significantly. So the changes in the official rate are mostly unanticipated by the market, and the surprise or announcement effect of the change is strong. This indicates that monetary policy was the leading variable in the period under consideration and that market rates followed the official rates. This result explains why long rates on the yield curve change less than proportionally after an official interest rate change, as market participants do not consider the official interest rate move, or the fundamental economic variable that causes the central bank to react, to be permanent. The limited reaction of long rates can also be due to the presence of future lower or higher short interest rate anticipations in the long rate before the official move actually took place. This effect, although present, was less strong in our results.

The same model was estimated for Germany. The degree of persistence is estimated at 0.89, but the coefficient is significantly less then one, implying that the underlying economic shocks are perceived to be mean reverting. In other words, German monetary policy is considered to be effective in targeting its objectives. The somewhat lower value for Belgium is explained by the higher volatility of short-term interest rates. But the differences in the formulation of policy targets between Belgium and Germany makes a comparison of these parameter values quite irrelevant.

The estimate of the surprise parameter θ is very low for Germany: the probability that the Bundesbank will not change its rate is estimated to be as low as 0.12% on average. Market participants are quite convinced that the Bundesbank will follow its normal reaction function; therefore they can almost fully anticipate official rate movements. The change in the official interest rate has only a weak surprise effect on the market rates at the day of change. The low correlation between the Repo rate and the money market rates (see Table 1) on the day of change is also found in Deutsche Bundesbank (1996). In this analysis it is mentioned that the low reaction of market rates is enhanced by the use of fixed-rate tenders, implying prior announcement of the rate of interest. Furthermore, changes in the repo rate are made in small and frequent steps, which may also contribute to a lower surprise effect. Although such technical reasons contribute to the low value of the estimated parameter, the obtained strong result does seem to confirm that the Bundesbank policy is perceived by market participants to be clear and credible. To that end the Bundesbank supplements its set of monetary policy instruments with an effective information strategy which makes its policy transparent and accountable (Schmid and Asche, (1997)). The German experience demonstrates that a credible monetary policy implies a relatively low reaction of long-term interest rates to official policy moves and that the latter may, for a substantial part, be anticipated by the market participants well in advance of official interest rate changes. The term structure of interest rates may, therefore, anticipate the future policy stance of the central bank. When analysing the impact of monetary policy one should therefore clearly distinguish that part which was already anticipated by the market.

Table 1

Impact of official interest rate changes on the market interest rate

OLS and multivariate non-linear LS estimates

Change in market interest rate: day after (*t*) – day before (*t*)	**Belgium[1]: change in central rate**		**Germany: change in Repo rate**	
Overnight	0.87	(0.07)	0.33	(0.16)
1-month eurorate	0.83	(0.12)	0.15	(0.07)
3-month eurorate	0.61	(0.10)	0.13	(0.06)
6-month eurorate	0.60	(0.09)	0.15	(0.06)
12-month eurorate	0.52	(0.06)	0.23	(0.06)
2-year domestic rate	0.26	(0.05)	0.15	(0.05)
5-year domestic rate[2]	0.13	(0.04)	0.06	(0.05)
10-year domestic rate			0.01	(0.05)
Change in market interest rate: day before (*t*) – day after previous change (*t*-1)				
Overnight	0.32	(0.14)	0.68	(0.17)
1-month eurorate	0.37	(0.12)	0.79	(0.11)
3-month eurorate	0.30	(0.11)	0.80	(0.10)
6-month eurorate	0.44	(0.11)	0.76	(0.10)
12-month eurorate	0.31	(0.11)	0.74	(0.12)
2-year domestic rate	0.20	(0.12)	0.44	(0.14)
5-year domestic rate[2]	0.20	(0.11)	0.25	(0.14)
10-year domestic rate			0.12	(0.13)
Model parameters				
θ : measure of surprise	0.84	(0.02)	0.12	(0.03)
ρ : measure of persistence	0.80	(0.07)	0.89	(0.05)

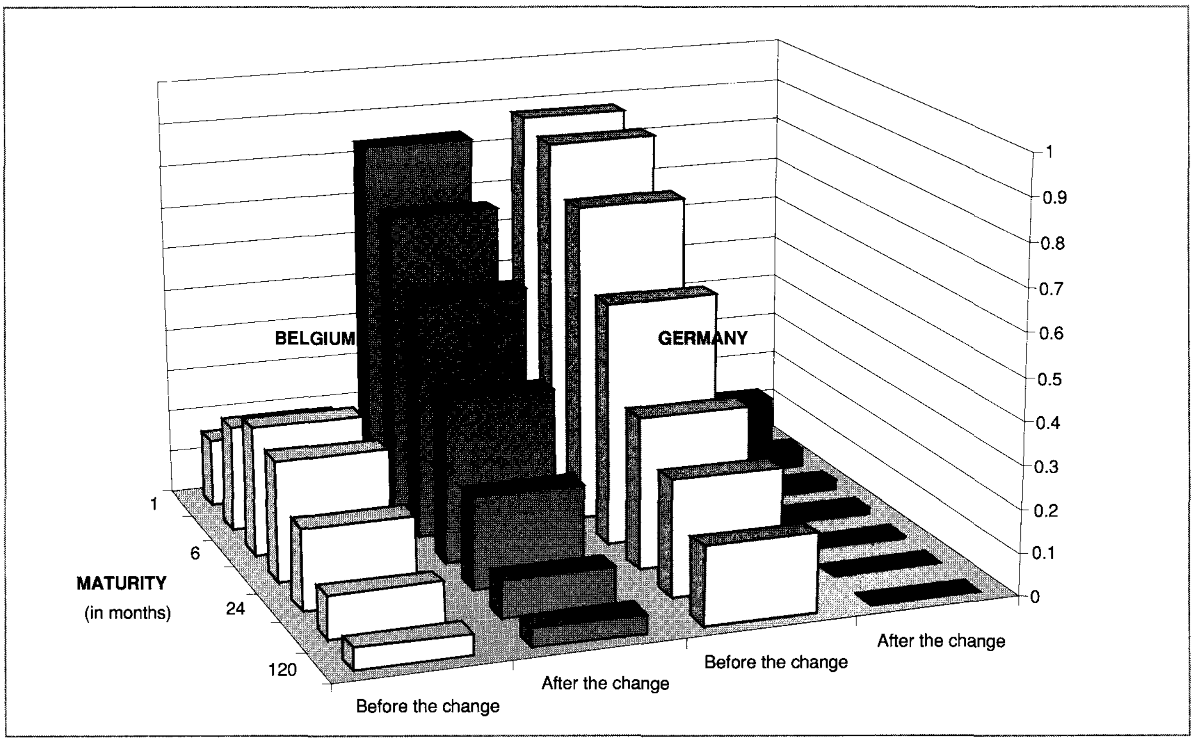
Note: Estimation period 31/1/1991-1/9/1997.

[1] Including 6 interest changes between 23/7/1993 and 15/9/1993. [2] Six years and more for Belgium.

Graph 1 differentiates surprise and anticipation effects of official interest rate changes. The higher surprise effect obtained for Belgium as compared to Germany is the result of differences in monetary policy objectives and instruments. The intermediate objective of monetary policy in Belgium is the stability of the DM/BF exchange rate. Exchange rate tensions, however, come quite suddenly and unpredictably, making the immediate response of the central bank equally sudden and non-anticipated in advance. Furthermore, in the absence of a system based on reserve requirements, banks may have less room to anticipate expected interest rate movements in Belgium.

Graph 1

Estimated term structure response to a one percentage point increase in the official rate



But, in any case, in both countries the central bank has a strong impact on the yield curve. This curve, therefore, reflects the "stance" of monetary policy as well as market expectations about future interest rates and economic conditions. This makes the interpretation of the yield curve more complicated. To derive useful information from it for monetary policy purposes, a correct interpretation of the underlying factors that explain the specific form of the yield curve at a given moment in time is needed. To that end we will try to identify the different underlying macroeconomic shocks that can be considered to be the main driving forces behind movements in interest rates, real growth and inflation. This should contribute to our understanding as to how these different shocks can explain the behaviour of the yield curve. But before doing this we first comment on the frequently observed correlations between the slope of the yield curve and future inflation and real GDP growth.

2. Correlations between term spread, growth and inflation

Monetary policy is basically forward looking. Since monetary policy actions affect the economy only with considerable lags, central banks need indicators of the future stance of macroeconomic variables such as inflation and real growth. According to the Consumption Capital Asset Pricing Model, a nominal interest rate on an asset with a given maturity should be related to the expected nominal growth rate over a time horizon that corresponds to the term to maturity of the asset considered. The difference between the yields on two assets with different maturities (that is the slope of the yield curve between two points) should therefore be correlated with the market participants' expected change in future nominal growth. The expected nominal growth rate can itself be decomposed into the expected real rate of growth and the expected inflation rate. We therefore analyse the correlation between the slope of the term structure involving assets with different maturities, on the one hand, and future changes in inflation and real economic growth on the other.

2.1 Correlations between the term structure and future inflation

We estimate the following equation:

$$\pi_t^j - \pi_t^k = \alpha + \beta\left(i_t^j - i_t^k\right)$$

where: π_t^j = average inflation between month t and j future months

π_t^k = average inflation between month t and k future months

i_t^j = interest rate in month t on an asset with j months of residual maturity

i_t^k = interest rate in month t on an asset with k months of residual maturity.

Table 2

Slope of the term structure and future inflation acceleration in Germany

***j* - *k* months**	α	β	R^2
6 - 3	0.04 (0.09)	-0.34 (0.47)	0.00
12 - 3	-0.11 (0.19)	-0.04 (0.12)	0.00
12 - 6	-0.07 (0.07)	0.09 (0.19)	0.00
24 - 12	-0.01 (0.14)	0.23 (0.22)	0.01
36 - 12	-0.14 (0.19)	0.47 (0.21)	0.07
48 - 12	-0.22 (0.29)	0.66 (0.22)	0.12
60 - 12	-0.42 (0.40)	0.87 (0.27)	0.18
72 - 12	-0.51 (0.51)	0.96 (0.30)	0.21
84 - 12	-0.66 (0.57)	1.02 (0.30)	0.24
96 - 12	-0.68 (0.61)	0.98 (0.29)	0.23
108 - 12	-0.64 (0.61)	0.94 (0.30)	0.23
120 - 12	-0.18 (0.66)	0.54 (0.28)	0.12

Notes: Estimation period 1972/I - 1996/I. Newey-West heteroskedasticity and autocorrelation-consistent standard errors for the OLS estimator are reported between brackets.

Table 3

Slope of the term structure and future inflation acceleration in Belgium

***j* - *k* months**	α	β	R^2
6 - 3	-0.03 (0.10)	-0.73 (0.47)	0.02
12 - 3	-0.05 (0.13)	-0.34 (0.27)	0.01
12 - 6	-0.04 (0.09)	-0.08 (0.34)	0.00
24 - 12	0.06 (0.10)	0.60 (0.14)	0.23
36 - 12	-0.02 (0.17)	0.80 (0.17)	0.25
48 - 12	0.03 (0.22)	0.74 (0.14)	0.21
60 - 12	-0.06 (0.27)	0.72 (0.14)	0.17

Notes: Estimation period 1977/III - 1996/II. Newey-West heteroskedasticity and autocorrelation-consistent standard errors for the OLS estimator are reported between brackets.

Tables 2 and 3 show positive correlations between inflation acceleration beyond one year in the future and interest rate differentials on assets with maturities exceeding 12 months in both Germany and Belgium. The term structure therefore does contain information on future inflation expectations (assuming that these expectations are formed rationally). This finding is in accordance with results generally found in the literature (see, for instance, Gerlach (1997) who reports results for Germany).

2.2 Correlations between the term structure and future real GDP growth

Theory predicts a positive correlation between the slope of the term structure and future real growth changes. In contradiction to this theoretical prediction, the literature rather reveals positive correlations between the slope of the term structure and future *levels* of real growth rates, instead of future *changes* in growth. We therefore estimated the following equation:

$$g_t^j = \alpha + \beta\left(i_t^j - i_t^k\right)$$

where: g_t^j = average real GDP growth between period t and j months into the future

i_t^j = period t interest rate on an asset with j months to maturity

i_t^k = period t interest rate on an asset with k months to maturity.

Table 4

Slope of the term structure and future real GDP growth in Germany

***j* - *k* months**	α	β	R^2
6 - 3	1.83 (0.36)	2.07 (1.18)	0.06
12 - 3	1.83 (0.35)	1.41 (0.48)	0.15
12 - 6	1.87 (0.31)	2.69 (0.77)	0.20
24 - 12	1.92 (0.40)	2.24 (0.58)	0.22
36 - 12	1.99 (0.29)	1.20 (0.32)	0.22
48 - 12	2.03 (0.21)	0.92 (0.21)	0.25
60 - 12	2.20 (0.19)	0.61 (0.17)	0.17
72 - 12	2.38 (0.18)	0.36 (0.15)	0.09
84 - 12	2.45 (0.19)	0.25 (0.14)	0.06
96 - 12	2.55 (0.19)	0.14 (0.11)	0.02
108 - 12	2.70 (0.21)	0.03 (0.12)	0.00
120 - 12	2.79 (0.16)	-0.06 (0.06)	0.00

Notes: Estimation period 1972/I - 1996/I. Newey-West heteroskedasticity and autocorrelation-consistent standard errors for the OLS estimator are reported between brackets.

The results for Germany and Belgium reported in Tables 4 and 5 show positive correlations between the slope of the term structure and future growth rates both in the short and medium-term segments of the term spread. Funke (1997) found the yield spread to be a useful indicator of the probability of future recessions in Germany.

Table 5

Slope of the term structure and future real GDP growth in Belgium

j - *k* months	α	β	R^2
6 - 3	1.87 (0.37)	0.50 (1.20)	0.00
12 - 3	1.89 (0.34)	0.76 (0.39)	0.05
12 - 6	1.89 (0.31)	1.94 (0.78)	0.11
24 - 12	2.30 (0.41)	1.66 (0.80)	0.17
36 - 12	1.98 (0.36)	1.43 (0.33)	0.26
48 - 12	1.86 (0.25)	1.33 (0.16)	0.52
60 - 12	2.07 (0.31)	0.42 (0.12)	0.03

Notes: Estimation period 1977/III - 1996/II. Newey-West heteroskedasticity and autocorrelation-consistent standard errors for the OLS estimator are reported between brackets.

3. Interpretation of the leading character of the term structure

The joint finding that the term spread is a leading indicator for future growth and inflation but that, at the same time, it can be strongly influenced by monetary policy actions is a quite relevant point when discussing the usefulness of the term spread as an indicator for monetary policy. Indeed, when the central bank interprets changes in the slope of the yield curve as an indicator of market expectations and acts on the basis of this indicator, but in doing so itself, affects the slope of the yield curve, policy instability may result.

It is therefore important to distinguish the shocks in the term structure that are induced by monetary policy actions themselves from the shocks that are generated by new and unanticipated events in the rest of the economy. If monetary policy actions are dominant, the term structure is primarily an indicator of the stance of monetary policy. It is only as far as the shocks elsewhere in the economy are relatively important that the term structure becomes a relevant information variable that should be taken into account when deciding on monetary policy actions. If that is the case, then a further step is required enabling the central bank to identify the underlying cause of the change in the slope of the yield curve, because the direction of a monetary policy reaction should depend on the nature of the particular shocks.

It is now generally accepted that the usefulness of the term spread as an information source for monetary policy heavily depends on the interpretation of the predictive power of the yield curve. In this respect, some authors have stressed the need for a structural interpretation of the term structure, and have pointed to the dangers that could result from a wrong interpretation of these signals (see, for example, Woodford (1994)). Different theoretical structural explanations are available to explain the predictive power of the term spread :

- the term structure can predict future interest rates, growth and inflation because it reacts nearly instantaneously to macro-economic shocks that drive the business cycle in the economy. The long rate can jump upwards if economic agents see the economy starting a growth phase that will lead to strong investment and credit demand and eventually, when full capacity is reached, will result in higher inflation. This reaction of the term structure to macro-economic shocks also anticipates the normal effect of the business cycle on monetary policy and on the financial behaviour of other sectors in the economy. Once capacity constraints start to increase and inflationary pressure builds up, monetary policy becomes more restrictive, flattening the yield curve. This will be followed by lower growth and inflation as the economy enters the downward phase of the cycle. In this channel, the observed correlation between the slope of the yield curve and future growth and inflation essentially reflects the reaction of monetary policy to movements in the business cycle;

- an alternative interpretation concentrates on the effects of monetary policy on the business cycle. Under this hypothesis the predictive power of the term spread reflects the impact of the short rate on future economic development. Both price rigidity in goods and labour markets and capital market imperfections (either in the form of imperfect substitution between financial assets and/or liabilities or in the form of imperfect distribution of liquidity over different sectors) explain why a monetary policy shock may have a transitory impact on real output and inflation. Under this hypothesis the information content of the spread about future inflation and growth, is due to lagged effects of innovations in the short-term interest rate;
- a third interpretation explains the "unique" predictive power of the term spread by the specific information that the spread contains on market expectations about future interest rates and the underlying determinants, like inflation and real growth. Under this hypothesis the information in the term structure is incorporated in the innovations in the long-term interest rate, which, under this hypothesis, is thought to be much more sensitive to expectations than short-term rates. Innovations in the long rate can reflect innovations in the expected real rate, in the expected inflation rate or in the risk premium:
- in a general equilibrium context, the consumption capital asset pricing model implies that economic agents determine the growth path of their consumption expenditures as a function of the real interest rate. A steep yield curve may therefore indicate an expected increase in the real growth rate of consumption over time;
- the long rate can also move in anticipation of an expected inflation shock. Such expectations can influence the behaviour of economic agents in the process of price and wage formation and create a self-fulfilling mechanism;
- finally, the long rate can increase when investors require a higher risk premium which can result both from an increase in the price of risk (higher degree of risk aversion) or from a deterioration in the perception of macro-economic fundamentals.

These alternatives but not necessarily mutually exclusive channels of the observed leading properties of the term spread are probably closely interrelated. The movements in the long-term rate can reflect "pure" shocks in the anticipation of economic agents but they may also be induced by adjustments to shocks occurring to other variables that drive the long rate. The same applies to the short-term rate: part of the movements in the short rate can be classified as the normal expected reaction of monetary policy to growth and inflation prospects, whereas the remaining part may be due to pure unexpected shocks, which may influence future business cycle movements.

3.1 Granger causality tests

As a first approach to distinguish these channels, we estimated dynamic equations for GDP growth and inflation in terms of own lags and lags of short and long-term interest rates, allowing us to perform Granger causality tests. Significant interest rate coefficients imply that the rates contain information that is not already available in the lags of inflation or growth. Also, the specific contribution of short-term versus long-term interest rates can be evaluated. Investigating the significance of each of these interest rates should indicate whether innovations in the short or the long rate or in both are the principal reasons behind the leading character of the spread.

Before discussing these results we should mention that in what follows we have always maintained the hypothesis that the variables used are stationary. For the German data this hypothesis is statistically confirmed, whereas Belgian inflation and interest rates are possibly non-stationary (Table 6). Repeating the causality tests using first-differences for potentially non-stationary series did not change any of the conclusions of the following tests. Therefore we do not report results with first differenced variables

Table 6

Unit root test (Augmented Dickey Fuller)

	Germany	**Belgium**	**Difference: Belgium – Germany**
GDP growth	-11.59**	-8.49**	-10.26**
CPI inflation	-3.02**	-2.02	-3.30**
Short-term interest rate	-3.75**	-2.29	-2.79*
Long-term interest rate	-2.97**	-1.57	-1.59
Term spread	-3.66**	-4.15**	-3.41**

Notes: Sample period: 1970:1 - 1997:1 – quarterly data (1972:1 - 1997:1 for Belgian GDP). **(*) indicates significance at the 5% (10%) level.

Table 7 reports the marginal significance levels of interest rates in the dynamic equations for GDP growth and inflation. For Belgium the results indicate that the short-term interest rate has a stronger predictive power as compared to the long-term interest rate, but both rates are hardly statistically significant in the GDP equation and not at all in the inflation equation. This result contrasts with the significant result in the simple expectation tests for both growth and inflation as discussed above. The diverging results can be explained by the different forecast horizon of both tests, different measures of the term spreads, and restrictions that are implied by empirically estimating the theoretical expectations hypothesis. The marginal influence or contribution of the domestic interest rate shocks to the real economy do not seem to be statistically very strong according to this test.

The results improve somewhat if we also introduce the German interest rates into the Belgian equations. Lags of the German short-term interest rate have a significant influence on both GDP and inflation in Belgium. Again the long rate performs less well, and only the domestic long rate is significant in the GDP equation. The finding that the German short-term interest rate tends to dominate the domestic interest rate, does not need to surprise. The domestic interest rates are "disturbed" by the short-term volatility of the exchange rate and the reaction of monetary policy to exchange rate shocks. As these disturbances have often been temporary in nature, they should not have had a strong impact on economic activity. This weak impact of domestic short interest rates is also confirmed by the results from the study of the transmission channel of monetary policy in Belgium (see Dombrecht and Wouters (1997)). As the German short-term interest rate is less disturbed by such short-term volatility effects, it is a better measure of the more permanent shocks that matter more for economic activity. The significance of the German interest rate may also capture an indirect effect. As the German short rate is important for German output and inflation, this will spill over to Belgian growth and inflation through bilateral trade. The effect of the German short rate on German growth and inflation is indeed confirmed by the data. The long-term interest rate is also less significant in Germany.

This first result seems to indicate that, to a large part, the often found leading character of the term spread disappears when past business cycle movements are included in the information set. Only the short-term interest rate seems to contain "exogenous" or innovating shocks that affect the future course of growth and inflation. This suggests that especially the monetary policy impact on future activity gives the interest spread its unique predictive character. Innovations in expectations that are present in the long-rate movements, do not significantly Granger cause future economic activity.

Table 7

Marginal significance levels of the short and long-term interest rates for forecasting GDP growth and inflation

Dependent variables	No. of lags	Short rate DM	Long rate DM	Short rate BF	Long rate BF	German GDP growth	German CPI inflation	Belgian GDP growth	Belgian CPI inflation
Germany									
GDP growth (quarter)	1 – 5	0.02	0.45			0.00	0.08		
CPI inflation (quarter)	1 – 5	0.05	0.16			0.26	0.00		
Short rate DM	1 – 5	0.00	0.14			0.48	0.14		
Long rate DM	1 – 5	0.97	0.00			0.95	0.10		
Belgium: domestic interest rates									
GDP growth (quarter)	1 – 4			0.06	0.10			0.39	0.85
CPI inflation (quarter)	1 – 4			0.18	0.78			0.37	0.00
Short rate BF	1 – 4			0.00	0.07			0.74	0.10
Long rate BF	1 – 4			0.74	0.00			0.90	0.28
Belgium: domestic and German interest rates									
GDP growth (quarter)	1 – 4	0.01	0.20	0.21	0.04			0.05	0.74
CPI inflation (quarter)	1 – 4	0.00	0.02	0.30	0.60			0.90	0.04
Short rate BF	1 – 4	0.02	0.99	0.04	0.31			0.82	0.37
Long rate BF	1 – 4	0.07	0.78	0.25	0.00			0.76	0.27

Notes: Sample period : 1970:1 - 1997:1 – quarterly data. For each forecasted variable, the entries give the marginal significance levels (p-values) for omitting all lags of the variable indicated in the column heading from an unrestricted OLS equation that also included a constant and *n* lags of growth and inflation. P-values not greater than 0.05 indicate statistical significance. The number of lags was selected, based on the absence of autocorrelation and the significance of the last lag starting from 12 lags.

The results also correspond with those of Estrella and Mishkin (1997). They find that for Germany the predictive power of interest rates is fully captured by the short-term interest rate and that the term spread as such does not contain extra information about economic growth. But they do not make a similar test for inflation. Their results for the United States, on the contrary, tend to indicate that the term spread may dominate the short-term rate as a predictive source for future growth. Contrary to this result, however, Bernanke and Blinder (1992) found that the Federal funds rate was the dominant variable, among other rates and money measures, in predicting economic activity. Smets and Tsatsaronis (1997), using a SVAR interpretation of the different types of shocks, conclude that inflation expectation shocks in long-term interest rate do not have an important real impact in Germany, while it cannot be neglected in the United States.

3.2 Cholesky decomposition of the forecast error variance

To investigate further the importance of the different shocks, and to distinguish endogenous from exogenous interest rate fluctuations, we estimate a VAR system for growth, inflation, short and long-term interest rates. A simple Cholesky decomposition of the error structure, where interest rates rank last, give a first indication of the importance of "pure" exogenous interest rate shocks in the total explanation of the growth and inflation process.

Consider the four-variable VAR system written in autoregressive form:

$B(L)\ x_t = e_t$ with covariance matrix Σ_e.

The Cholesky factorisation decomposes the covariance matrix in terms of a lower triangular matrix θ_o: $\Sigma_e = \theta_o * \theta'_o$. The reduced form errors e_t can then be rewritten in terms of four orthogonal innovations u_t with $e_t = \theta_o u_t$.

The Cholesky decomposition of the covariance matrix implies that the errors of the reduced-form process are restructured into four independent shocks. By setting GDP first in the equation order, the error term in the GDP equation is the first independent real shock, that affects the variables coming behind in a specific order (first inflation, then interest rates). The remaining error in the inflation process is the second independent shock process. By setting interest rates last, we actually deduct from the reduced-form errors, the information already present in the macro-economic variables. This corrects for simultaneity in the error shocks of the different variables and allows us to simulate exogenous shocks in the rates of interest, i.e. interest rate shocks that are independent of those affecting growth and inflation. This further allows a separation of the normal endogenous reaction of interest rates, either through the reaction function of the central bank or through the reaction of private agents to macro-economic disturbances, from pure exogenous shocks in monetary policy or in the long-term yield. The first component of the interest rate process results from the operation of non-monetary shocks, while only the remaining disturbances are considered as of a specific monetary nature. In this respect it, can be noted that the responses of the variable ranking last are independent of the ranking of the preceding variables. We set the long-term interest rate in the last row of the VAR system, behind the short rate, as it is more likely that the short rate does not react contemporaneously to long-rate innovations than vice versa.

Table 8 gives the contribution of each of the four shocks, identified according to the Cholesky decomposition to the explanations of the four variables under consideration. They result from a four-order VAR system estimated on quarterly data for Belgium over the period 1972:1 to 1997:1. The contribution of the two exogenous interest rate innovations to GDP-growth and inflation is very limited. For Belgium, the two interest rate shocks explain no more than 5% of the variability of growth and inflation. For Germany we estimated a similar four-order VAR system over the period 1970:1 to 1997:1 and found the contribution of the short rate to real growth rate fluctuations to be somewhat higher (around 10%). The contribution of both interest rate shocks does not exceed 15% for growth and even less for inflation.

Table 8

Forecast error variance decomposition: Cholesky decomposition

Order of the variables: GDP growth, inflation, short rate, long rate

	Belgium							
Quarters	**GDP growth**				**Short rate**			
	Output	Inflation	Short rate	Long rate	Output	Inflation	Short rate	Long rate
0	1.00	0.00	0.00	0.00	0.13	0.02	0.84	0.00
4	0.90	0.07	0.03	0.01	0.14	0.23	0.62	0.01
8	0.89	0.07	0.03	0.01	0.13	0.25	0.58	0.04
12	0.89	0.07	0.03	0.01	0.11	0.28	0.56	0.05
16	0.89	0.07	0.03	0.01	0.11	0.29	0.54	0.06
20	0.89	0.07	0.03	0.01	0.10	0.30	0.53	0.06
24	0.89	0.07	0.03	0.01	0.10	0.31	0.53	0.06
	Inflation				**Long rate**			
	Output	Inflation	Short rate	Long rate	Output	Inflation	Short rate	Long rate
0	0.00	1.00	0.00	0.00	0.00	0.01	0.17	0.82
4	0.05	0.91	0.02	0.03	0.02	0.13	0.17	0.68
8	0.05	0.90	0.03	0.03	0.05	0.27	0.20	0.47
12	0.05	0.90	0.03	0.03	0.05	0.32	0.24	0.39
16	0.05	0.89	0.03	0.03	0.05	0.33	0.26	0.36
20	0.05	0.89	0.03	0.03	0.05	0.34	0.27	0.34
24	0.05	0.89	0.03	0.03	0.05	0.34	0.28	0.33
	Germany							
Quarters	**GDP growth**				**Short rate**			
	Output	Inflation	Short rate	Long rate	Output	Inflation	Short rate	Long rate
0	1.00	0.00	0.00	0.00	0.00	0.08	0.92	0.00
4	0.81	0.09	0.07	0.03	0.07	0.13	0.79	0.01
8	0.78	0.10	0.09	0.04	0.14	0.21	0.63	0.01
12	0.77	0.11	0.09	0.04	0.14	0.22	0.60	0.03
16	0.76	0.11	0.09	0.04	0.14	0.22	0.59	0.06
20	0.76	0.11	0.09	0.04	0.14	0.22	0.58	0.06
24	0.76	0.11	0.10	0.04	0.14	0.23	0.58	0.06
	Inflation				**Long rate**			
	Output	Inflation	Short rate	Long rate	Output	Inflation	Short rate	Long rate
0	0.00	1.00	0.00	0.00	0.13	0.05	0.19	0.64
4	0.05	0.85	0.04	0.06	0.15	0.23	0.17	0.45
8	0.06	0.84	0.05	0.05	0.14	0.40	0.12	0.34
12	0.06	0.83	0.06	0.05	0.12	0.45	0.13	0.30
16	0.06	0.82	0.06	0.05	0.12	0.45	0.14	0.29
20	0.07	0.82	0.06	0.05	0.12	0.45	0.14	0.29
24	0.07	0.82	0.06	0.05	0.12	0.45	0.14	0.28

These results may, however, underestimate the importance of interest rate shocks since the application of the Cholesky decomposition method implies independence of the inflation process from monetary policy innovations (the implied independence of the GDP growth process from short-term interest rate innovations is in this respect less restrictive as it is normally assumed that monetary policy does not affect growth immediately). To get a more meaningful interpretation of the different shocks it is necessary to use a structural approach that allows the imposition of theoretical restrictions to identify the independent structural innovations that drive the joint VAR system.

3.3 A structural-VAR approach

We will not discuss the underlying theoretical model in detail here but limit ourselves to a short discussion of the theoretical restrictions that are actually used. For a discussion of the theoretical model we refer to Fuhrer and Moore (1995), and Smets and Tsatsaronis (1997). Following the literature on Structural VAR models, we distinguish supply, demand, monetary policy and long term interest rate shocks by the following assumptions:

- only supply shocks have long run real effects. This assumption follows from a theoretical model with a vertical long-run supply curve. It implies three zero-restrictions on the long-term impact of the three remaining shocks on GDP growth. A positive supply shock (for instance, a productivity increase or a real wage decrease) should have a negative impact effect on inflation and, depending on the reaction function of the monetary authorities, on the short term interest rate;
- demand shocks can influence all four variables on impact: an increase in demand in a model with price rigidity affects both real GDP and inflation positively. Again depending on the reaction function of the monetary authorities, the short interest rate will increase and, depending on the persistence of the short rate increase, the long rate will follow. In the long run, demand shocks should be neutral for output;
- monetary policy does not affect real growth contemporaneously: this is a restriction on the short-term impact. A restrictive monetary policy should be reflected in a higher short-term and long-term interest rate, with probably a negative impact effect on inflation. Long-run neutrality of monetary policy on output is a generally accepted restriction;
- the fourth type of shock can be identified as an innovation in the risk premium or inflationary expectations that drive the long-term interest rate. This shock is identified by two short-term restrictions: it should not influence contemporaneously real growth and monetary policy actions as measured by the short-term interest rate. On the contrary an increase in the long-term interest rate, may be accompanied by an increase in inflation as far as it reflects an "inflation scare" effect that is subsequently realised or self-fulfilling.

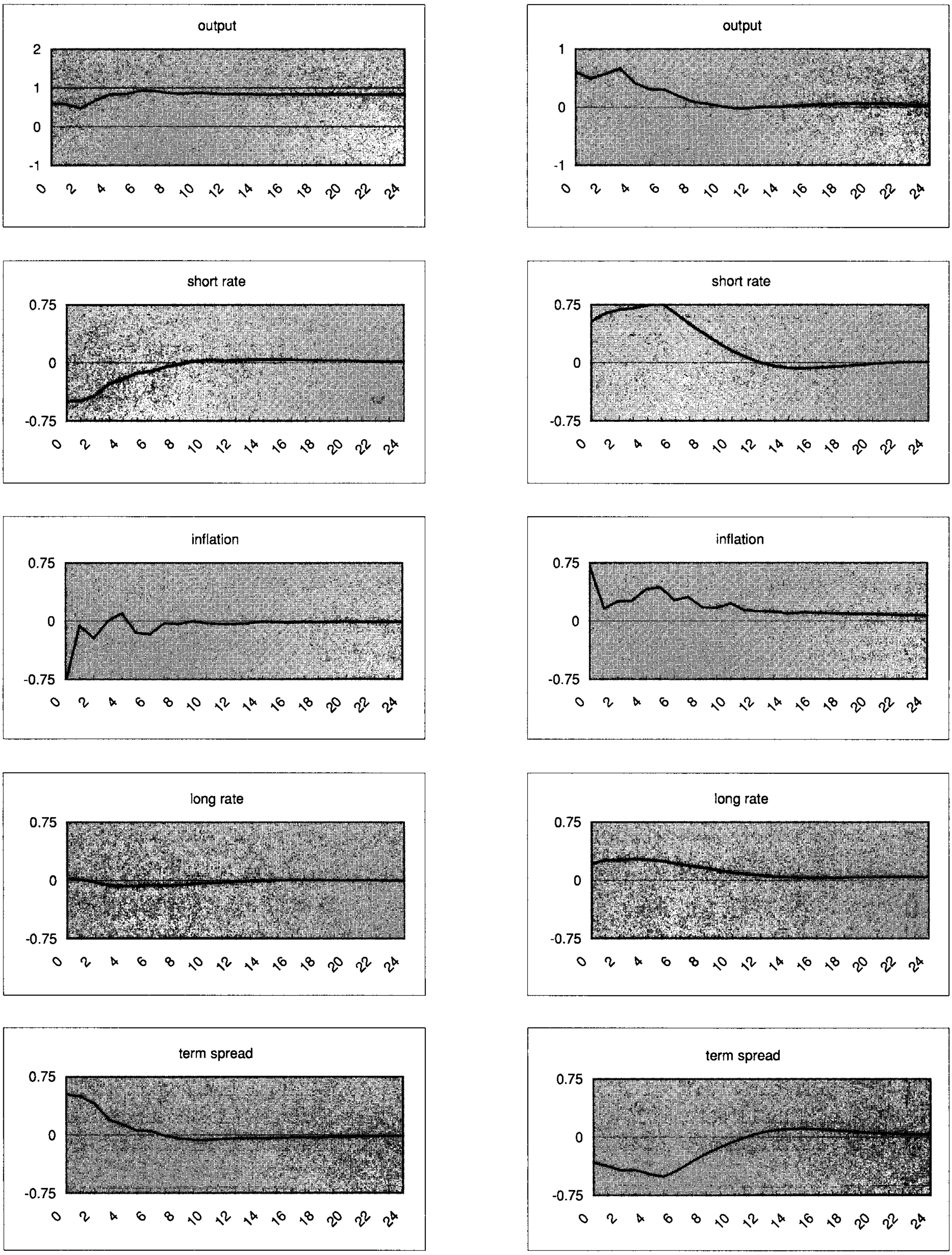
These theoretical considerations provide six restrictions on the parameters of the decomposition matrix θ_o: three short-term restrictions directly on the elements of θ_o, and three long-term restrictions on $\theta(1)$ with $\theta(1) = C(1)*\theta_o$ and $C(1)$ the sum of the vector moving average coefficients of the model $x_t = B(L)^{-1} * e_t = C(L) * e_t$. So, together with the ten restrictions on θ_o, resulting from the symmetry of the covariance matrix, the identification of the model is complete.

This structural model, applied to the VAR system for the German data, produces expected results. The impact of the different shocks is shown in the impulse responses of the four types of shocks, and these correspond to the theoretical expectations (Graph 2). Using the theoretical restrictions to identify the monetary policy shock and the shocks in the long-term interest rate, the contribution of the interest rate shocks in the variance decomposition of inflation is strongly increased (Table 9). According to our estimation results, the autonomous shocks in monetary policy explain about 40% of the variability of inflation and the contribution of the long-term interest rate shocks is increased towards 20%. The contribution to real growth fluctuations remains limited to a joint 12%. These results are close to the ones obtained by Smets and Tsatsaronis (1996), although our long-term interest rate effect is somewhat stronger.

Graph 2a

Impulse response of the SVAR for Germany
Quarterly basis

Suppy shock | Demand shock

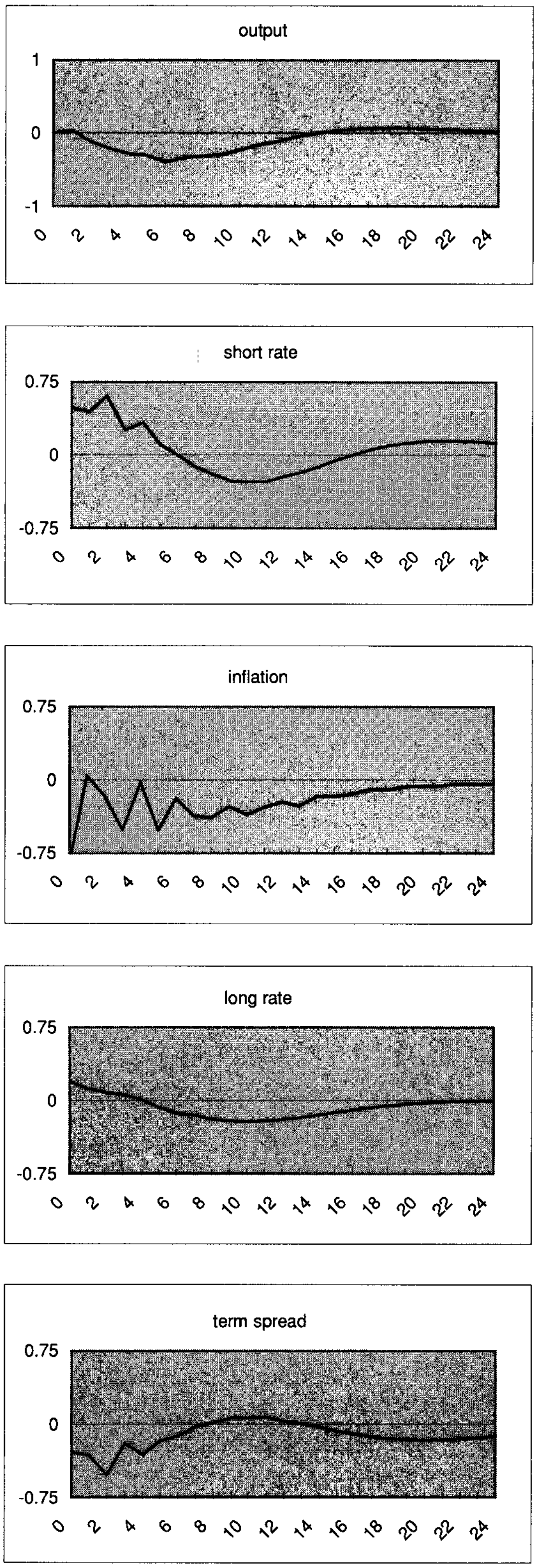


Graph 2b

Impulse response of the SVAR for Germany

Quarterly basis

Monetary policy shock



Long interest rate shock

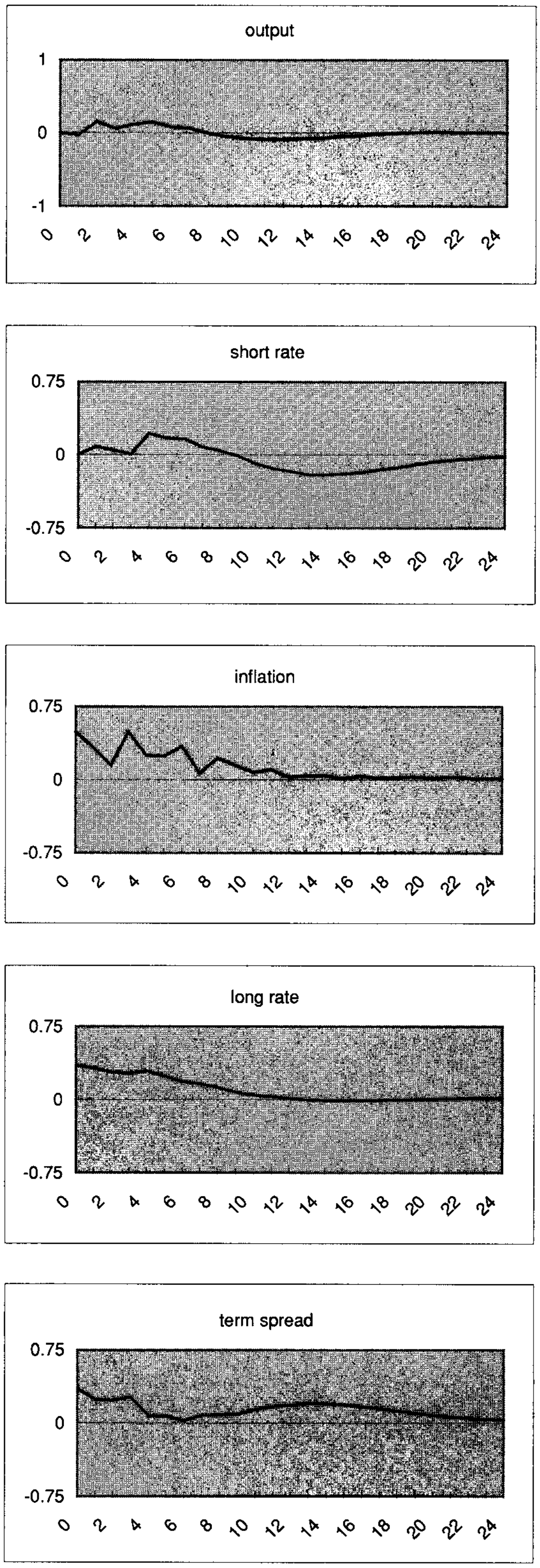


Table 9

Forecast error variance decomposition: SVAR model

Belgium								
Quarters	**GDP growth**				**Short rate**			
	Supply	Demand	Mon. pol.	Long rate	Supply	Demand	Mon. pol.	Long rate
0	0.72	0.28	0.00	0.00	0.01	0.28	0.71	0.00
4	0.61	0.33	0.05	0.01	0.06	0.47	0.46	0.01
8	0.61	0.33	0.05	0.01	0.06	0.48	0.42	0.04
12	0.61	0.33	0.05	0.01	0.06	0.49	0.39	0.06
16	0.61	0.33	0.05	0.01	0.06	0.50	0.37	0.07
20	0.61	0.33	0.05	0.01	0.06	0.51	0.36	0.07
24	0.61	0.33	0.05	0.01	0.06	0.52	0.35	0.07
	Inflation				**Long rate**			
	Supply	Demand	Mon. pol.	Long rate	Supply	Demand	Mon. pol.	Long rate
0	0.26	0.67	0.07	0.00	0.01	0.02	0.15	0.82
4	0.22	0.67	0.09	0.02	0.02	0.17	0.11	0.70
8	0.23	0.66	0.09	0.03	0.02	0.38	0.10	0.49
12	0.22	0.66	0.09	0.03	0.03	0.44	0.12	0.41
16	0.22	0.66	0.09	0.03	0.04	0.46	0.13	0.38
20	0.22	0.66	0.09	0.03	0.04	0.47	0.13	0.36
24	0.22	0.66	0.09	0.03	0.04	0.48	0.14	0.35
Germany								
Quarters	**GDP growth**				**Short rate**			
	Supply	Demand	Mon. pol.	Long rate	Supply	Demand	Mon. pol.	Long rate
0	0.49	0.51	0.00	0.00	0.32	0.37	0.30	0.00
4	0.44	0.47	0.04	0.05	0.20	0.55	0.24	0.01
8	0.42	0.47	0.05	0.06	0.15	0.64	0.19	0.02
12	0.42	0.47	0.06	0.06	0.14	0.61	0.22	0.03
16	0.41	0.46	0.07	0.06	0.13	0.59	0.22	0.05
20	0.41	0.46	0.07	0.06	0.13	0.58	0.22	0.06
24	0.41	0.46	0.07	0.06	0.13	0.58	0.23	0.06
	Inflation				**Long rate**			
	Supply	Demand	Mon. pol.	Long rate	Supply	Demand	Mon. pol.	Long rate
0	0.30	0.24	0.34	0.11	0.00	0.21	0.19	0.60
4	0.22	0.26	0.32	0.21	0.01	0.37	0.08	0.54
8	0.17	0.27	0.36	0.20	0.02	0.39	0.11	0.48
12	0.15	0.27	0.39	0.19	0.02	0.35	0.21	0.41
16	0.15	0.26	0.41	0.18	0.02	0.34	0.25	0.39
20	0.15	0.27	0.41	0.18	0.02	0.34	0.25	0.39
24	0.15	0.27	0.41	0.18	0.02	0.34	0.25	0.38

When applying the model to Belgium, we introduce the German data (contemporaneously and one lag) into the equations as representatives of exogenous foreign shocks. These exogenous variables explain already a large part of both macro-economic variables and interest rates. What is left are domestic shocks, which can be identified with the same combination of short and long-term restrictions as those mentioned above. The results, summarised in the impulse responses in Graph 3, are again acceptable:

- the supply shock has a positive effect on growth and a negative one on inflation. Both interest rates decline only gradually but very persistently. The strong negative impact effect on the short-term interest rate, as found in Germany, cannot be expected in Belgium where monetary policy is fully concentrated on the exchange rate objective;
- the demand shock has a strong positive impact on both real growth, inflation and interest rates. Here, the impact effect on the short-term interest rate is acceptable as a positive demand shock, probably originating in public expenditures and causing a worsening of the current account. This is likely to have a direct impact on the exchange rate and therefore, given the objective of monetary policy, on the short-term interest rate;
- monetary policy shocks, reflected in a strong short-term interest rate increase, only have a small negative impact effect on inflation and are followed by a moderate increase in the long rate. The small effect on inflation is acceptable as the short-term interest rate in Belgium is mainly increased to offset exchange rate pressure. As this policy had to gain credibility during the estimation period, the small positive long-term effect on inflation can be considered as acceptable;
- the long-term interest rate effect is not followed by a similar movement in inflation but is rather reflected in the short-term interest rate. This result seems logical given that such shocks in Belgium were more likely related to a decrease in the risk premium on Belgian franc investments, following perceived improvements in the underlying fundamentals (government deficit and current account). A shock in the long-term Belgian franc rate should therefore be considered as a risk premium shock rather than as an innovation in inflation expectations.

From the variance decomposition in Table 9, it follows that the explanatory power of domestic monetary policy and long-term interest rate innovations in the Belgian case, remains very limited as far as real growth and inflation is concerned. However the German interest rates, that enter the equations as exogenous variables, do have a significant impact on real growth and inflation. The contribution of these shocks can, however, only be evaluated if a joint VAR system for the two countries has been estimated.

The decomposition of interest rates, growth and inflation into the four underlying structural shocks that drive the economy, now allows us to look for the reasons of the observed leading character of the term structure with respect to growth and inflation. This information can be obtained from the decomposition of the covariance between growth and inflation with the lagged term structure.

In Germany, the positive covariance between growth and the lagged term structure is mainly explained by supply and demand shocks (Graph 4). But monetary policy shocks also contribute to the positive covariance. By pushing up the short rate and flattening the yield curve, a restrictive policy has a negative effect on growth in the following quarters. The positive correlation between the lagged term spread and growth is especially strong in the short run, up to one year. This result corresponds to the expectations tests in Section 2.2 where the predictive content of the term structure, as far as growth was concerned, was mainly found in the short end of the maturity structure.

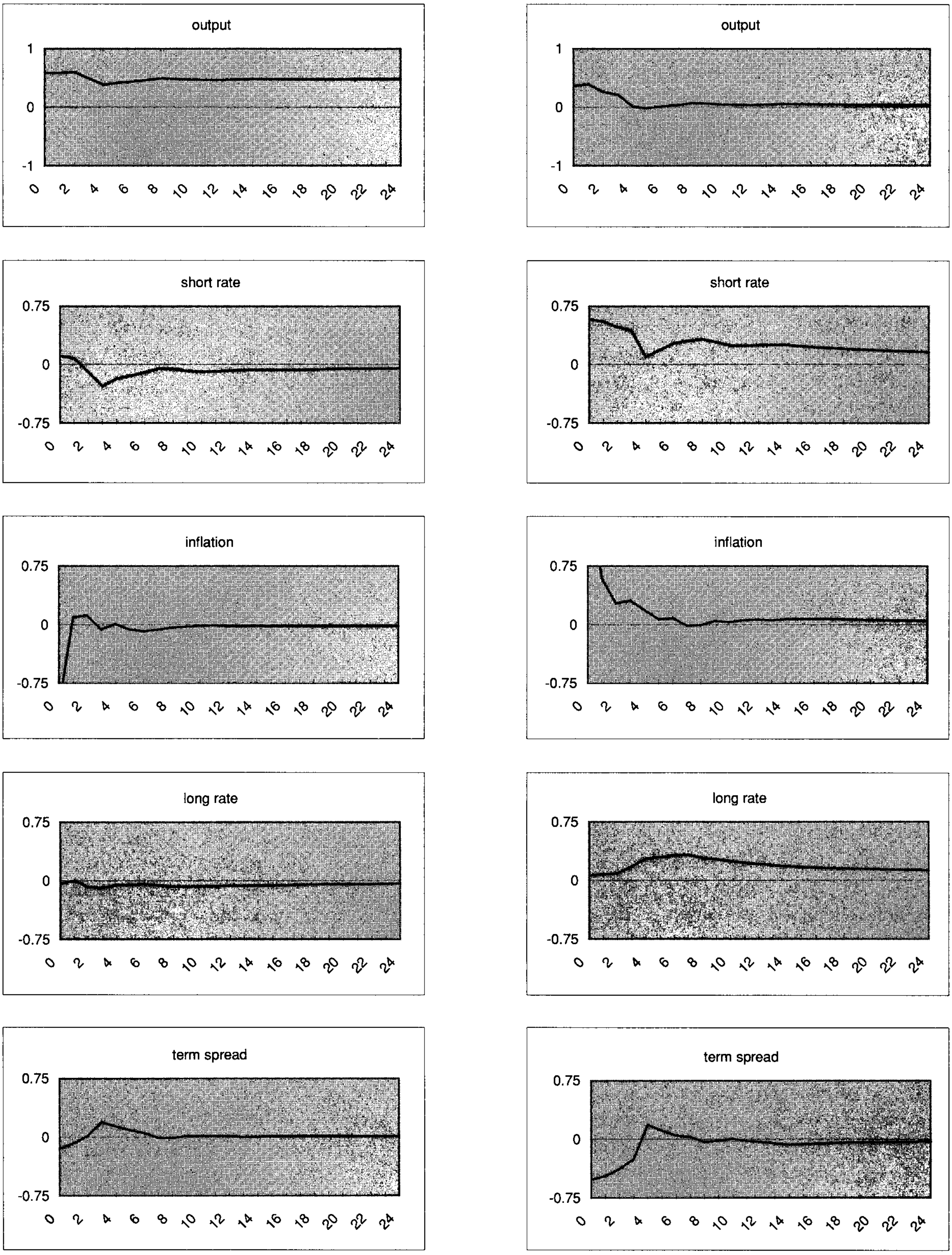
A similar analysis decomposing the covariance between the lagged term spread and the rate of inflation is presented in Graph 4. The covariance is close to zero up to one year and becomes slightly positive only for longer lags. This net result is obtained as a sum of significant individual contributions that tend to work in opposite directions. On the one hand, supply and especially demand shocks suggest a strong negative correlation: a demand shock has a positive effect on inflation but is followed by a strong decrease in the term spread as monetary policy reacts restrictively such that the term spread tends to be negatively correlated with current and future inflation. On the other hand, monetary shocks and innovations in inflationary expectations or in the risk premium induce positive correlation between the term spread and future inflation. At longer lags these positive contributions tend to dominate the negative ones induced by demand and supply shocks. The results provide an explanation for the conclusion reached in many tests of the information content of the term structure

Graph 3a

Impulse response of the SVAR for Belgium
Quarterly basis

Suppy shock

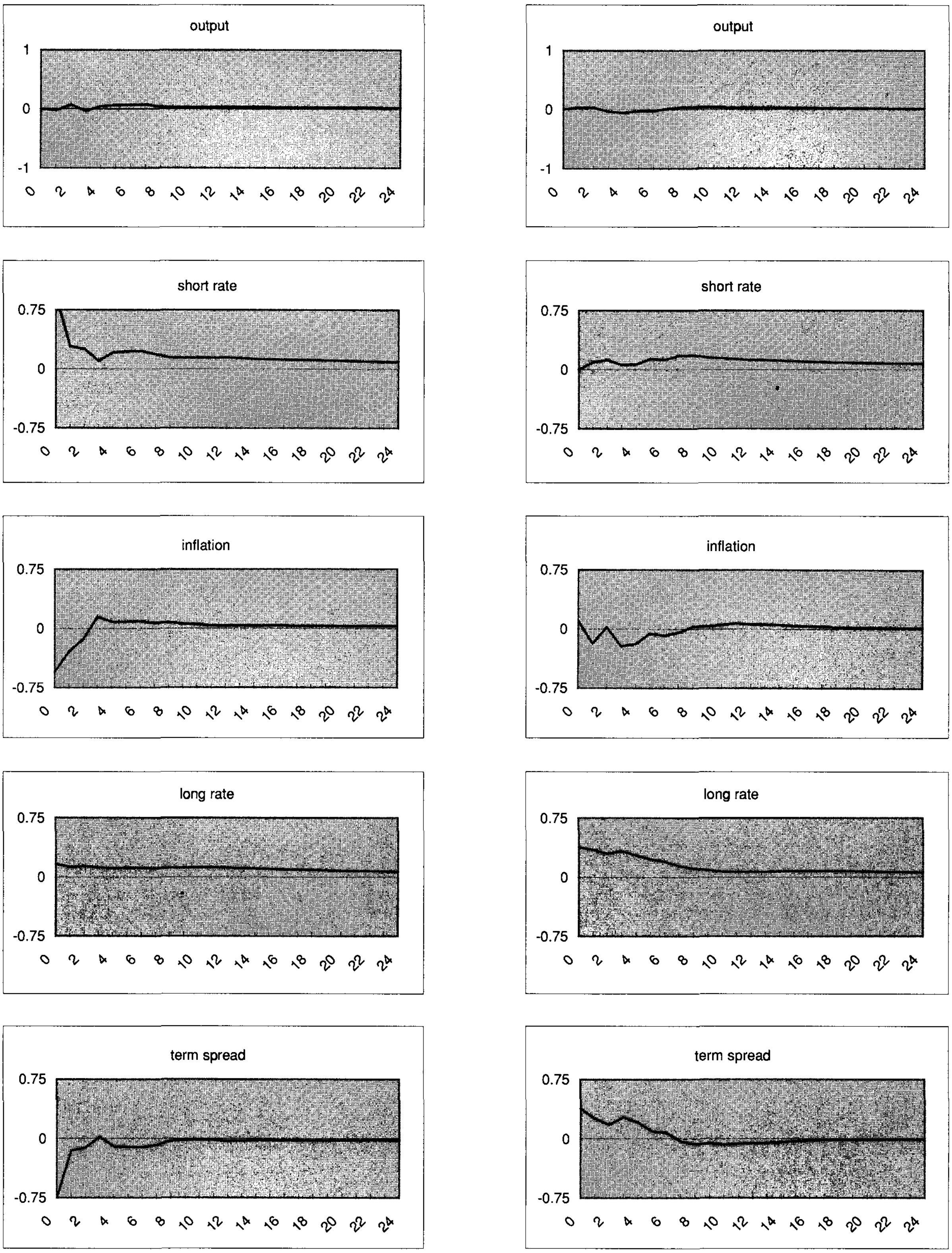
Demand shock



Graph 3b

Impulse response of the SVAR for Belgium
Quarterly basis

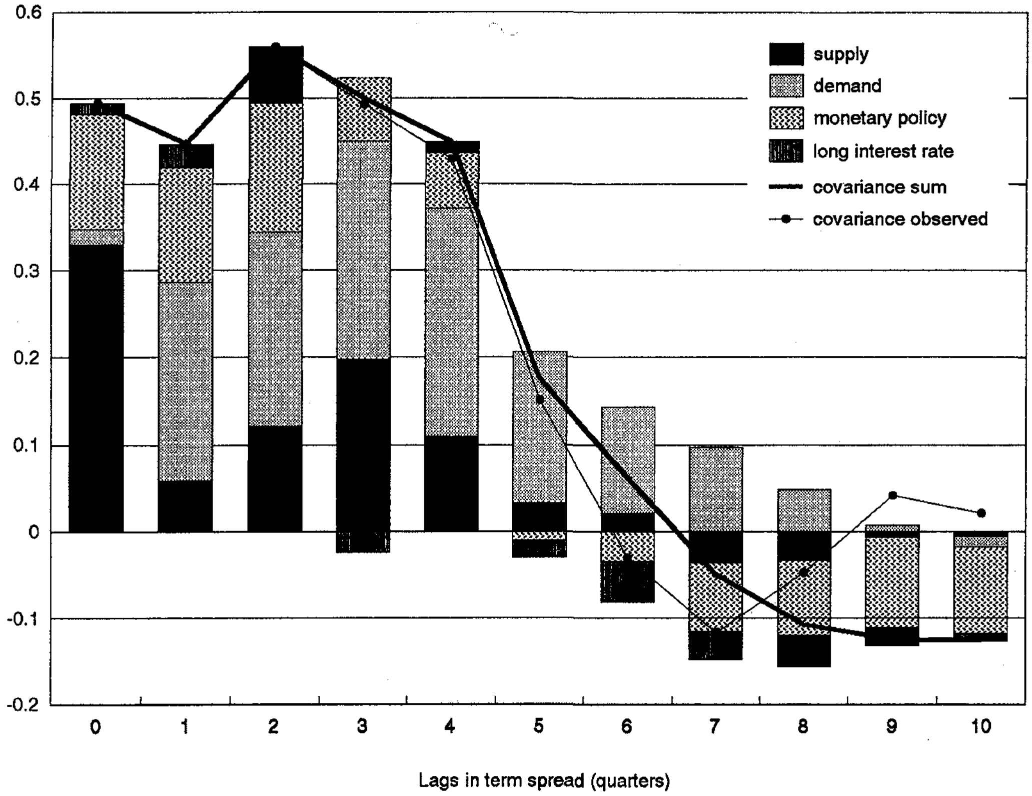
Monetary policy shock | Long interest rate shock



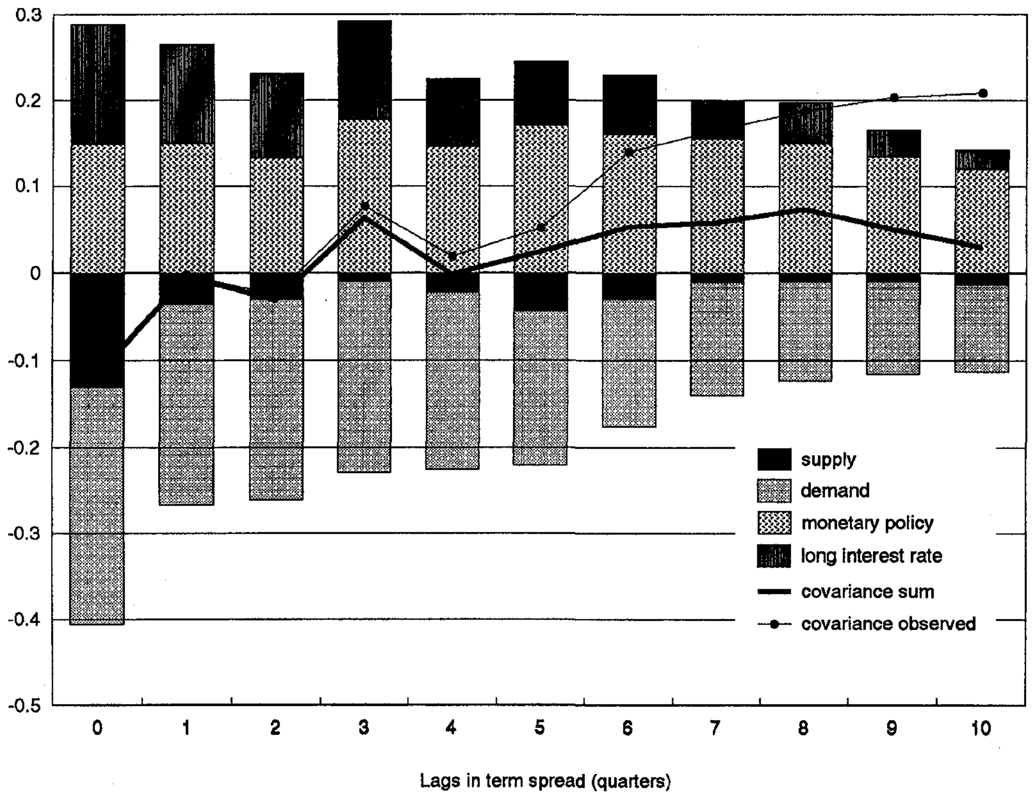
Graph 4

Structural decomposition of growth and inflation/term spread covariance in Germany

Growth



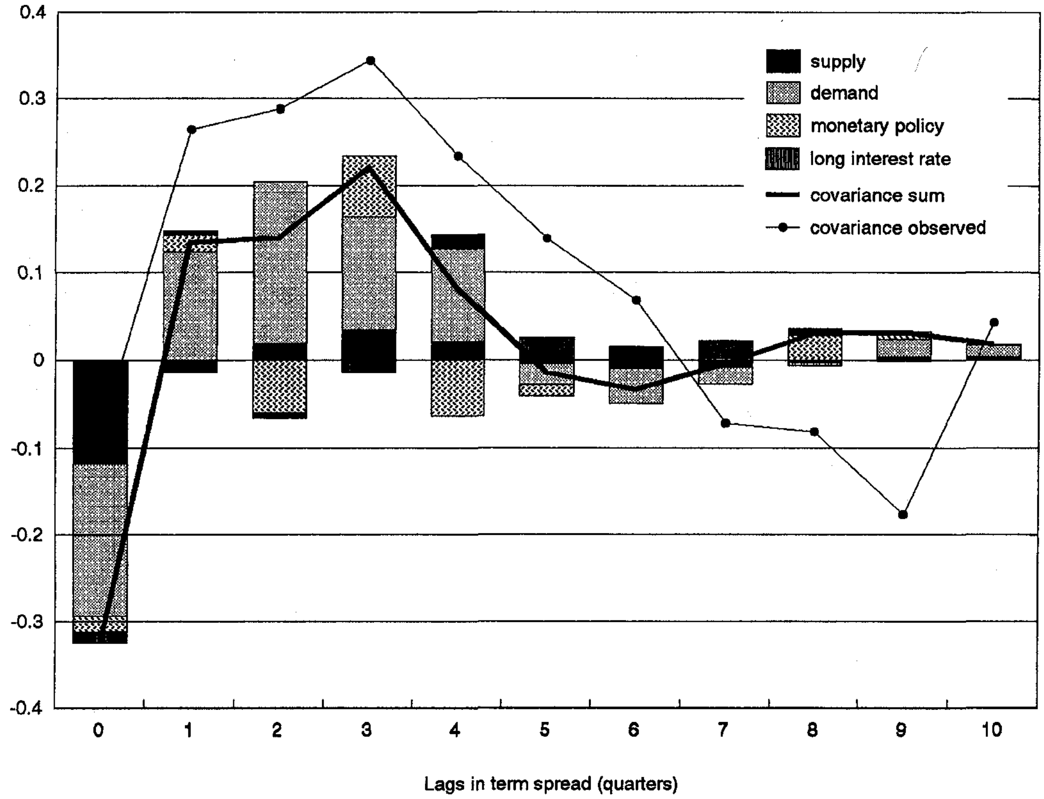
Inflation



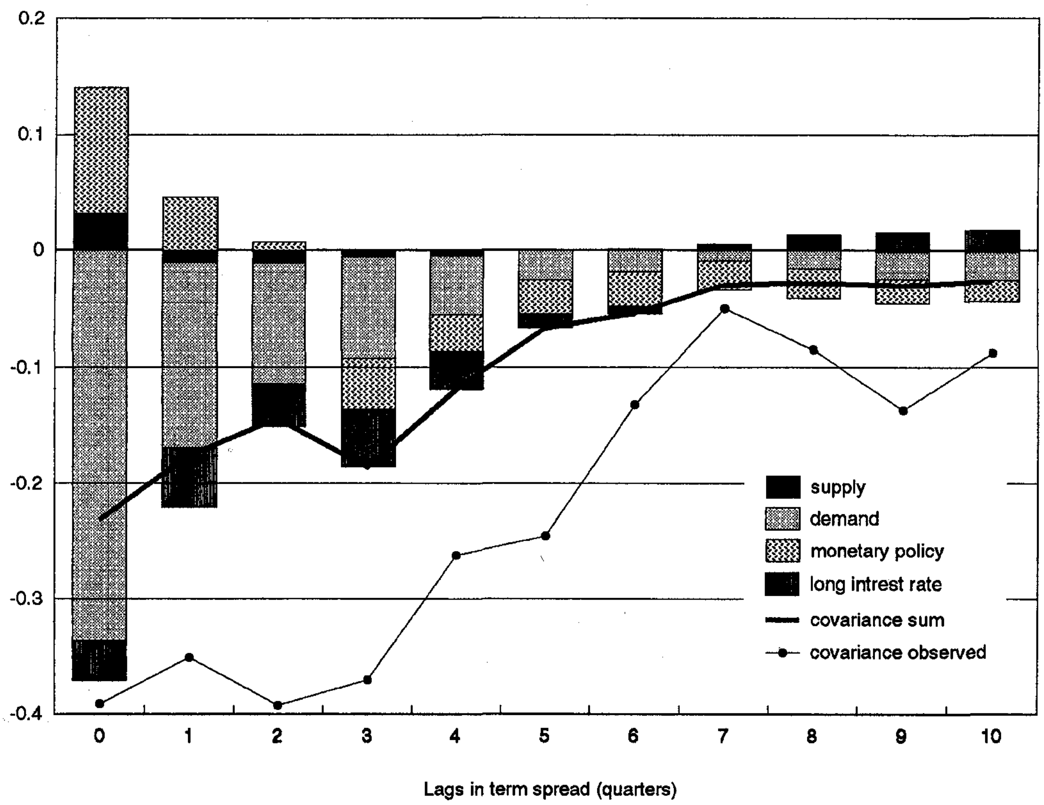
Graph 5

Structural decomposition of growth and inflation/term spread covariance in Belgium

Growth



Inflation



according to which the term spread is a predictor of future inflation at longer time horizons only (see also Section 2.1). Graph 4 also indicates that our VAR model is not able to fully reproduce the observed positive covariances between inflation and the term structure especially at longer lags. An insufficient number of lags in the estimated VAR model is the main reason for this imperfection.

Applying the same exercise to Belgian data (Graph 5) gives rise to several observations. The exogenous German variables in the model provide a strong contribution to the explanation of the observed covariances (this is demonstrated by the significant deviation between the observed and explained values of the covariances). As far as the domestic shocks are concerned, only demand surprises contribute significantly to the covariances (positive and negative contributions to the covariance between the term spread and, respectively, growth and inflation). As explained above, neither domestic monetary policy shocks nor the unexpected movements in the risk premium present in the long-term rate interest rate are able to offset the contribution of demand shocks to the observed negative correlation between the term spread and future inflation in Belgium.

In general, the results of this SVAR exercise correspond with previous tests for Belgium, indicating that domestic interest rate shocks are less important in explaining the correlation between the term spread and future growth and inflation rates, because these comovements tend to be caused mainly by shocks originating in Germany.

When applying these kinds of models to monetary policy analysis, it should be remembered that their results only provide average outcomes conditioned by the sample period. Specific questions related to actual monetary policy issues or to the interpretation of the actual slope of the yield curve, need to based on a careful analysis of the actual economic situation in terms of the shocks discussed in this paper.

Conclusions

Traditional tests of the information content of the term structure of interest rates reveal that the slope of the yield curve is correlated with future changes in inflation and with future real growth. It was therefore concluded that the yield curve's inclination contains information on market participants' expectations concerning the future course of inflation and growth. On the other hand, evidence has also emerged indicating that the slope of the yield curve may be significantly affected by actual or by market perceived changes in official central banks' interest rates. The joint events whereby the term spread reflects market participants' expectations about inflation and that, at the same time, the term spread is itself affected by central banks' reactions, may imply policy instability whenever the central bank effectively tries to use the term spread as an indicator of future inflation in setting its official interest rates. It is therefore important to analyse the relative importance of the contribution of monetary policy shocks as compared to other types of innovations that may affect the slope of the yield curve. If monetary shocks were found to be dominant, then the term spread would mainly reflect the stance of monetary policy.

Granger causality tests indicate that short-term interest rates have a much stronger predictive impact on future inflation and growth as compared to long rates, suggesting that the observed positive correlations between the term spread and future output growth and inflation is mainly due to monetary policy reactions. For example, a rise in the official interest rate, as a reaction to unfavourable inflation signals, lifts the short end of the yield curve. Its impact on the long end of the term structure may be relatively minor in so far as inflation shocks are perceived as persistent but essentially temporary in nature. Market participants therefore expect the short rate to come back to a "normal" level in the long run in line with the central bank's inflation target, such that long-term interest rates are hardly affected by a temporary rise in short-term interest rates.

Other evidence can be found from a structural vector autoregression model. Such a VAR system in inflation, growth, short and long-term interest rates was used to identify supply, demand, monetary policy and long-term interest rate shocks. We decomposed the observed positive covariance between the slope of the term structure and future inflation and found that, especially at the longer end of the yield curve, unexpected shocks in short-term interest rates explain a large part of the observed positive covariance. Innovations in the long-term interest rate, e.g. due to new inflation signals, only explain a minor part of this covariance.

These results imply that prudence is required when using the slope of the yield curve as an indicator of (forward looking) monetary policy formulation. Changes in the shape of the yield curve may signal different types of unexpected events in both the real and financial sectors of the economy. Those changes should therefore first be carefully interpreted in terms of underlying structural shocks in order to identify the information which such changes are actually signalling.

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Stock market returns, inflation and monetary regimes

Giuseppe Grande, Alberto Locarno and Massimo Massa*

Introduction

Since Fisher's initial contribution in the early thirties, several studies have looked at the ability of different assets to provide a hedge against inflation.[1] The Fisher hypothesis, relying on the idea that the monetary and real sectors of the economy are largely independent, states that expected asset returns should move one-to-one with expected inflation. In principle this hypothesis is applicable to any instrument that can serve to transfer wealth through time, but it should especially apply to assets representing physical capital, such as real estate and shares in the capital of a company. These assets should also provide a hedge against unexpected inflation. However, empirical studies have often concluded that the Fisher hypothesis is not well supported by the data; more surprisingly, its failure appears clearest for equities.

Theoretical as well as applied research has shown that the relation between stock prices and inflation is influenced by economic policy, and by monetary policy in particular. This paper focuses on the relation between stock returns, inflation and monetary policy. The working hypothesis is that the market interprets inflation differently according to a latent variable that captures the effects of shifts in the stance and the credibility of monetary policy, as well as those of changes in the institutional framework in which the central bank operates. Financial markets react differently to inflation news, depending on the monetary policy regime they perceive to be the prevailing one. When the central bank is thought to be strongly committed to price stability, even a small surge in inflation expectations induces the market to fear a strong monetary policy reaction, which would lead to higher interest rates, lower economic activity and lower expected dividends. As a consequence, stock prices drop, and the negative relationship between stock returns and expected inflation usually found in the literature obtains. This is essentially the so-called proxy hypothesis proposed by Fama (1981) and developed by subsequent studies, as will be explained in the next section.

The empirical framework adopted in this paper – applied to data on the Italian stock exchange covering the last twenty years – relies upon the present-value relation of Campbell and Shiller, and makes use of a Markov-switching model to identify regimes associated with different policy environments. The analysis focuses on the inflation information contained in stock returns, and does not address the issue of the possible effects of equity prices on real activity.[2]

After presenting a brief review of the main arguments put forward to explain the failure of the Fisher hypothesis in stock markets, we provide an initial assessment of the relation between asset returns and inflation in Italy in the second section. Then we present the VAR model with Markov switching and the decomposition of the βs of a portfolio according to the present-value relation. The last section discusses the methodological issues raised in the paper and sets out the main conclusions.

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1 See Fama and Schwert (1977) and the survey by Rovelli (1984).

2 A careful comparison of the balance sheets of household and enterprise sectors in the major industrial countries can be found in Kneeshaw (1995).

1. Different explanations of the relation between stock returns and inflation

Various explanations of the failure of the Fisher hypothesis when applied to the stock market try to interpret empirical results in terms of spurious correlations and omitted variables. Some studies have also addressed the issue on theoretical grounds.

Fama (1981) argued that the sign on inflation is due to the fact that inflation acts as a proxy for omitted variables. Given that high inflation anticipates low growth and that there is a positive relationship between expected economic growth and stock prices, there should be a negative relationship between inflation and stock prices. According to Stulz (1986), an increase in expected inflation leads to a fall in the real wealth of households, which in turn lowers the real interest rate and the expected return on the market portfolio. Geske and Roll (1983) relate the high rates of inflation during recessions to counter-cyclical monetary policy actions. The central bank responds counter-cyclically to real activity shocks: a drop in real activity leads to a higher public deficit which, in turn, induces an increase in money growth to the extent that the debt is monetised. An unanticipated drop in stock prices signals this chain of events, with the counter-cyclical expansion of the money supply reinforcing the "proxy" mechanisms proposed by Fama.

The perception of a clear link between stock prices and monetary and fiscal policy induced Kaul (1987) to focus on the relationship between monetary regimes and the Fisher equation. In particular, he showed how the counter-cyclical monetary policy regime in the post-war period generated a strong negative relationship between stock returns and changes in expected inflation; conversely, the relationship was positive under pro-cyclical monetary policy regime in the thirties. Furthermore, Kaul (1990) found evidence that the negative relation between stock returns and changes in expected inflation in the post-war years is particularly strong during interest rate regimes. More recently, Balduzzi (1993) proposed a VAR decomposition that reinterprets the proxy hypothesis, showing that both inflation and stock returns tend to anticipate future interest rate changes, albeit in opposite directions. Groenwald et al. (1997) examine the matter within the framework of a small macroeconomic model and find that the reduced form for the interest rate equation is much more complex than that used by Fama and Schwert and requires a larger set of variables to be explicitly taken into account. Though they propose and estimate a more refined specification, they find that the negative sign of the correlation coefficient survives the extension to the full model.

Söderlind (1997) uses a modified version of a model by Fuhrer and Moore (1995) to show that the sign and size of the correlation between stock returns and inflation in a closed economy depend on the objective function of the central bank. Suppose that (i) inflation is persistent but can be controlled via a Phillips effect; (ii) output is negatively related to real interest rates through an IS-type relation; (iii) the interest rate is set by the monetary authorities; and (iv) there are exogenous inflation shocks. Under these assumptions, if the central bank wants to stabilise output, it will move the nominal rate so as to keep the real interest rate constant: the nominal interest rate then entirely reflects changes in expected inflation and the Fisher effect is complete. If the central bank targets inflation instead, it will use the nominal rate in order to allow the real rate to move as much as is required to stabilise expected and actual inflation. If this policy is successful, the nominal interest rate will be mainly correlated with the real rate and the Fisher relation will not be satisfied.

Focusing on econometric issues, Evans and Lewis (1995) reformulate the Fisher puzzle in terms of a time-varying model. They do not search for an economic rationale for the failure of the Fisher hypothesis for bond rates, but try to explain it in terms of small sample biases induced by the infrequency of shifts in the inflation process during the post-war period.

In sum, previous literature has pointed out that contemporaneous regressions of stock returns on inflation expectations, while simple and useful, do not shed light on the channels through which macroeconomic news affects asset prices. Moreover, the co-movements of inflation and stock prices are clearly influenced by monetary policy and, more generally, by the policy environment.

Concerning the first issue, the proxy hypothesis put forward by Fama can be interpreted as an attempt to remove the influence of future output growth; similarly Geske and Roll try to neutralise the effects of monetary policy by including money supply as an additional explanatory variable in simple regressions of stock returns on expected inflation. More generally, Groenewold et al. stress that once we interpret the Fisher relation as the reduced-form equation of a macroeconomic model, we must allow for a large number of additional variables affecting stock returns in addition to the rate of inflation, namely the exchange rate and government consumption. Regarding the second issue, Kaul (1987, 1990) acknowledges that the correlation between stock returns and inflation is altered by policy actions and suggests dividing the sample period according to shifts in the policy regime in order to allow a proper evaluation of the Fisher effect. Finally, Söderlind claims that "the Fisher effect [...] is probably not carved in stone, but is likely to depend on monetary policy". As mentioned above, if the central bank wants to stabilise output, movements in nominal rates will parallel movements in inflation, while if it aims to preserve price stability the yield curve will not provide meaningful information about inflation expectations.

A clear example of the shortcomings of reduced-form models of stock returns is provided by Campbell and Ammer (1993). They cite the case of the reaction of the stock market to news about industrial production. This association could reflect either a link with changing expectations about future cash flows or some correlation with movements in future discount rates, perhaps because both industrial production and stock prices respond to interest rate changes. The only way to distinguish these channels is to deal explicitly with the relations linking stock prices to future dividends and required returns. This is the approach adopted in this paper.

2. Asset returns and inflation: a first step in the empirical analysis

As a first step in the empirical analysis, we replicate the approach developed by Fama and Schwert to draw a general picture of the relationship between asset returns and inflation in Italy.

The Fisher theory of interest assumes that the monetary and real sectors of the economy are largely independent. Expected real returns are uncorrelated with expected inflation, being determined by non-financial factors such as productivity of capital, time preferences and risk aversion: expected asset returns therefore move one-to-one with expected inflation. However, in order to assess whether financial assets provide a hedge against inflation, it is also necessary to analyse how nominal returns react to unexpected inflation.

To address these issues in a consistent framework, Fama and Schwert begin with the following equation (see equation (3) in Fama and Schwert (1977)):

$$E(R_{jt} \mid \phi_{t-1}) = E(i_{jt} \mid \phi_{t-1}) + E(\pi_t \mid \phi_{t-1}) + \gamma_j [\pi_t - E(\pi_t \mid \phi_{t-1})] \tag{2.1}$$

where R_{jt} is the nominal return on asset j from time t-1 to time t, ϕ_{t-1} is the information set at t-1, π_t is the inflation rate from time t-1 to time t and i_{jt} is the equilibrium real return.

On the basis of equation (2.1) and having a measure of the expected inflation rate, $E(\pi_t \mid \phi_{t-1})$, tests of the joint hypothesis that markets are efficient[3] and expected real returns and inflation are uncorrelated can be obtained from the following regression model:

$$R_{jt} = \alpha_j + \beta_j \pi_t^e + \gamma_j (\pi_t - \pi_t^e) + \varepsilon_{jt} \tag{2.2}$$

[3] That is, agents' expectations are the best possible assessment of the expected value of random variables given available information.

where, for simplicity, $\pi_t^e \equiv E(\pi_t \mid \phi_{t-1})$. If the coefficient β is not significantly different from 1, then the Fisher hypothesis cannot be rejected and the asset provides a complete hedge against *expected* inflation; if $\gamma = 1$, then the asset is a complete hedge against *unexpected* inflation; finally, if both β and γ are not significantly different from 1, then the asset is a complete hedge against *actual* inflation, and ex-post real returns and inflation are uncorrelated.

Fama and Schwert point out that the relation between nominal returns and unexpected inflation is not the same for all assets: while it is generally believed that real estate, common stocks and human capital[4] are hedges against both anticipated and unanticipated inflation, short-term securities, with fixed nominal payments, are entirely exposed to nominal shocks.

Fama and Schwert estimate the regression (2.2) on monthly US data for the period between January 1953 and July 1971. Building on previous work by Fama, the return on Treasury bills with a residual maturity of one month is used as a proxy for the expected value of inflation. They find that: (i) Treasury bills and bonds provide a hedge against expected inflation; (ii) private residential real estate hedges against expected as well as unexpected inflation; and (iii) labour income shows a weak correlation with inflation. The most striking result is obtained for common stocks, whose nominal returns appear to be negatively related to expected and, probably, unexpected inflation.

A crucial role in this kind of test is played by the measurement of expected inflation. Santoni and Moehring (1994) claim that the puzzle shown by Fama for stock returns can be accounted for once inflation expectations are properly measured. Three proxies for expected inflation have been used in the literature:

1. the nominal return on Treasury bills (Fama and Schwert, Mishkin (1990), and Kaul (1990));[5]
2. survey data on inflation expectations (Bomberger and Frazer (1981));[6]
3. expected inflation defined on the basis of a set of previously specified variables. Balduzzi (1993), for example, explicitly defines expected inflation by inverting a rational-expectations version of the standard quantity theory equation.

We apply the approach suggested by Fama and Schwert to Italian data on five different assets:[7] 3, 6 and 12-month Treasury bills; Treasury bonds;[8] and the value-weighted Milan stock exchange index. With regard to inflation expectations, since none of the aforementioned approaches is without shortcomings or is uniformly superior to the others, we try different alternatives. We use both the Forum-ME survey data and the fitted values of the projection of inflation on its own lags and the percentage changes in the exchange rate and industrial production; an additional attempt is made on

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4 They adopt the rate of change in per capita labour income as a proxy for the nominal return on human capital (this does not account for changes in capital values). If real labour income is to be independent of the price level, the measure must reflect inflation rate movements.

5 This choice is based on two hypotheses: (i) the expected real return on the short-term bill is constant through time and (ii) the market is efficient, so that the nominal return on the bill is equal to the constant expected real return plus the expected inflation rate; that is, it coincides with expected inflation apart from a constant factor.

6 Some shortcomings are inherent in the use of survey data, the main one being that the sample may not be representative of the whole economy. Also, it is certainly true that economists have better theories of how people take actions than they do of how they answer questions on surveys. Finally, unlike in a market where the participants back up their statements with money, it is less clear what it means when someone just expresses opinions about inflation or other variables.

7 The data used in the paper are described in Appendix 1.

8 Average yield on Treasury bonds with at least one year to maturity.

quarterly data by using the 3-month Treasury bill rate. The period covered runs from February 1979 to May 1997 for monthly data and from the second quarter of 1979 to the first quarter of 1997 for quarterly data.

The results for monthly data are reported in Table 1a. For each asset, two regressions are estimated according to the proxy considered for expected inflation. The estimated values of the parameters and their standard errors are shown in the first three columns, while the probabilities that β and γ are equal to one are shown in the last two columns.

The hypothesis that these assets are hedges against expected inflation is rejected with respect to Treasury bills and, to a lesser extent, Treasury bonds, but the estimated value of the parameter β is positive and markedly different from zero, suggesting that the assets provide at least a partial hedge against expected inflation. Conversely, the parameters associated with unexpected inflation are never significantly different from zero.

It is very important to observe that the test on ß is actually a joint test of three hypotheses: lack of correlation between the expected values of the real rate and inflation; market efficiency; and the Fisher hypothesis. The above results must be interpreted with caution, because the rejection of the null hypothesis could be due to the fact that a fully developed market for government securities in Italy emerged only at the end of the eighties.[9]

A further warning is due because the effect of taxes on capital income has not been taken into account. Since financial assets are usually taxed, a change in inflation that is fully transmitted to nominal interest rates does not leave the lender with the same pre-inflationary real return: nominal returns have to move more than proportionately to leave the after-tax real rate unaffected. A proper treatment of this issue, which is complicated by the fact that tax incidence is not the same for all investors and assets, is beyond the scope of this study. Note, however, that the coefficient on inflation in equation (2.2) has to be greater than one if after-tax returns are to provide a complete hedge against inflation.

When the inflation forecast is measured by survey data or by the fitted values of a time-series model, the results obtained with quarterly data are very close to those found with monthly figures (Table 1b). When possible, the regressions on quarterly data also make use of the 3-month Treasury bill as a proxy for expected inflation. In this case, the hypothesis of a complete hedge against expected inflation cannot be rejected at standard confidence levels, while the estimated value for γ is still not consistent with the hypothesis of perfect coverage against unexpected inflation.

As regards stocks, in all models neither β nor γ are significant at standard confidence levels and the proportion of the variance of stock returns explained by the regression expectations is very low (about 2%). However, the estimated effect of inflation forecasts on stock returns is positive, as expected. This is an important difference with the results obtained on US data with similar methodologies.

To check for instability in the coefficients and to see how the estimated relation between nominal returns and expected inflation has moved through time, rolling regressions on a ten-year window, spanning the whole sample period, have been run on quarterly data; stability analysis has only been applied to the regressions that use the inflation forecasts of the Forum-ME survey. The estimated values for β and its confidence bands are plotted in Figure 1; the horizontal dashed line

9 A screen-based secondary market for government securities was introduced in May 1988 and grew quickly. The volume of transactions in Treasury bills on the secondary market has always been very thin. For this reason, the returns on Treasury bills used in the paper are those determined through competitive auctions on the primary market. It must be noted that until March 1989 the Treasury set a floor for the bid price, which often turned out to be binding; this constraint lessened the link between the average yield at auction and agents' expectations. In March 1989 the lower bound for bids was removed for all maturities; Grande (1994) provides evidence that the ability of the primary Treasury bills market to signal agents' expectations improved since that date.

indicates the points where the parameter is equal to 1: when the line is inside the confidence band, the hypothesis of the asset being a complete hedge against expected inflation cannot be rejected. This appears to be true for government securities since the eighties. However, the estimated β varies considerably over the period, and its standard error clearly shows a tendency to widen. The rolling estimates of the β parameter for stocks confirm the failure of this simple test of the Fisher hypothesis for the Italian stock exchange.[10]

Table 1a

Effects of expected and unexpected inflation on asset returns in Italy

Expected inflation proxy	α	β	γ	R^2	σ	H_0: β=1 complete hedge against expected inflation	H_0: γ=1 complete hedge against unexpected inflation
			(a) 3-month Treasury bills				
Forum-ME survey	0.006 (0.0003)	0.639 (0.0357)	-0.045 (0.0335)	0.608	0.002	0.00	–
AR model	0.007 (0.0002)	0.417 (0.0254)	-0.072 (0.0409)	0.552	0.002	0.00	–
			(b) 6-month Treasury bills				
Forum-ME survey	0.006 (0.0003)	0.637 (0.0367)	-0.041 (0.0344)	0.594	0.002	0.00	–
AR model	0.007 (0.0002)	0.420 (0.0257)	-0.077 (0.0414)	0.551	0.002	0.00	–
			(c) 12-month Treasury bills				
Forum-ME survey	0.006 (0.0002)	0.616 (0.0354)	-0.033 (0.0332)	0.596	0.002	0.00	–
AR model	0.007 (0.0002)	0.407 (0.0249)	-0.062 (0.0401)	0.549	0.002	0.00	–
			(d) Treasury bonds				
Forum-ME survey	0.007 (0.0002)	0.586 (0.0364)	-0.024 (0.0341)	0.561	0.002	0.00	–
AR model	0.008 (0.0002)	0.393 (0.0251)	-0.060 (0.0404)	0.528	0.002	0.00	–
			(e) Stocks				
Forum-ME survey	-0.003 (0.0113)	2.748 (1.576)	-0.130 (1.477)	0.006	0.069	–	–
AR model	0.001 (0.0085)	2.093 (1.046)	-0.966 (1.684)	0.011	0.069	–	–

Notes: Equation (2.2) is run on monthly data for the period 1979:2-1997:5. The statistic R^2 is adjusted for the degrees of freedom. σ is the standard error of the regression. Numbers in parenthesis are parameter standard errors. The last two columns show the probabilities of being wrong in rejecting the indicated hypotheses; they are reported only for those cases in which the estimated parameter is different from zero at a 5% confidence level. A description of the data is given in the Appendix.

[10] For almost the whole sample period, the hypothesis that the value of the parameter is equal to zero cannot be rejected.

All in all, the results in Table 1 confirm the findings in Fama and Schwert, though there is evidence that the relation between stock returns and expected inflation is positive in Italy as expected. These results signal that the Fisher hypothesis is not well supported by the empirical evidence, especially for stock returns.

In the rest of the paper we will try to model the relation between inflation and stock returns more accurately, taking into account the role played by policy regimes.

Table 1b

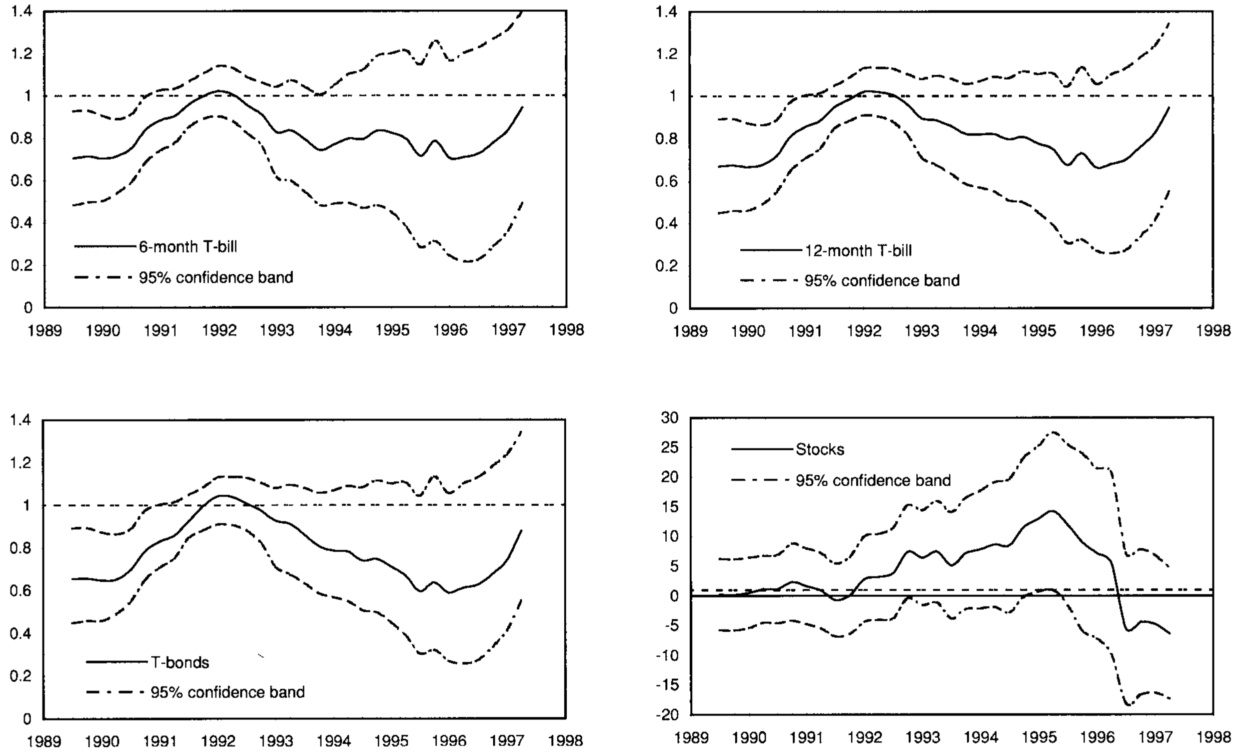
Effects of expected and unexpected inflation on asset returns in Italy

Expected inflation proxy	α	β	γ	R^2	σ	H_0: β=1 complete hedge against expected inflation	H_0: γ=1 complete hedge against unexpected inflation
			(a) 6-month Treasury bills				
3-month Treasury bill	0.003 (0.0016)	0.944 (0.0487)	0.141 (0.0307)	0.847	0.003	25.71	0
Forum-ME survey	0.018 (0.0016)	0.689 (0.0757)	-0.124 (0.1042)	0.576	0.005	1.06	–
AR model	0.022 (0.0012)	0.458 (0.0480)	-0.102 (0.1060)	0.561	0.005	0.00	–
			(b) 12-month Treasury bills				
3-month Treasury bill	0.004 (0.0014)	0.921 (0.0419)	0.135 (0.0264)	0.877	0.003	6.23	0
Forum-ME survey	0.018 (0.0015)	0.661 (0.0719)	-0.108 (0.0989)	0.583	0.005	0.12	–
AR model	0.022 (0.0011)	0.441 (0.0458)	-0.081 (0.1011)	0.565	0.005	0.00	–
			(c) Treasury bonds				
3-month Treasury bill	0.005 (0.0014)	0.887 (0.0426)	0.124 (0.0269)	0.864	0.003	1.03	0
Forum-ME survey	0.020 (0.0015)	0.614 (0.0726)	-0.082 (0.0999)	0.547	0.005	0.00	–
AR model	0.024 (0.0011)	0.418 (0.0458)	-0.066 (0.1011)	0.537	0.005	0.00	–
			(d) Stocks				
3-month Treasury bill	0.032 (0.0678)	1.194 (2.034)	2.248 (1.281)	0.019	0.128	–	–
Forum-ME survey	0.015 (0.0396)	1.427 (1.900)	2.754 (2.615)	0.018	0.128	–	–
AR model	-0.002 (1.310)	2.262 (1.181)	0.356 (2.618)	0.023	0.127	–	–

Notes: Equation (2.2) is run on quarterly data for the period 1979:II-1997:I. The statistic R^2 is adjusted for the degrees of freedom. σ is the standard error of the regression. Numbers in parenthesis are parameter standard errors. The last two columns show the probabilities of being wrong in rejecting the indicated hypotheses; they are reported only for those cases in which the estimated parameter is different from zero at a 5% confidence level. A description of the data is given in the Appendix.

Figure 1

Assets as hedges against expected inflation



Notes: Rolling regressions on a 10-year window running from 1979:II-1989:I to 1987:II-1997:I. The diagrams show the parameter in equation (2.2) associated with expected inflation. The proxy used for the latter is the inflation forecast of the Forum-ME survey.

3. Stock returns, inflation and monetary regimes in Italy

The model developed in this section builds on two considerations. First, the framework suggested by Fama and Schwert is not adequate for testing the Fisher effect. Being a restricted version of a reduced-form model, it does not provide any guidance on the selection of the relevant variables and runs the risk of identifying spurious correlations. Second, being dependent on the reaction function of the central bank, equation (2.2) is subject to structural instability. The literature surveyed in Section 1 largely supports these two claims.

The analysis is carried out by splitting the return on a stock or portfolio into two components: the riskless rate, proxied by the interest rate on 3-month Treasury bills, and the excess return. To explain excess returns we rely on the CAPM, while we use the present value relation along the lines suggested by Campbell and Shiller (1988a, b) to detect the channels through which macroeconomic factors affect βs and the market risk premia. This model has the advantage of relying on a sound theoretical basis, because it relates asset prices to their fundamental components. In particular, the decomposition by Campbell and Shiller allows us to express the innovation in the excess return of the stock market as a function of revisions in the expectations on the future values of dividends, excess returns, real interest rates and inflation.

In present value models a crucial role is played by assumptions about the way in which market participants forecast these fundamental variables. We assume that market participants approximate the evolution through time of the relevant state variables by means of a VAR process.

The effects of policy actions are accounted for by allowing the response of financial markets to news to depend on their perception of how the central bank responds to shocks to the economy. Unlike most studies, we do not explicitly define the monetary policy regimes themselves, but rather we try to infer them from market behaviour assuming that regime shifts are governed by an unobserved Markov process. That is, unlike Kaul (1990), we do not explicitly divide the sample period according to the monetary regimes, but rather model the latter as a latent variable in a Markov-switching model, thus allowing the data to speak for themselves. As long as we are able to approximate the way in which financial markets process information, we should succeed in providing a reasonable account of market expectations about policy actions. This is a standing feature of the paper for at least two reasons: first, it enables us to avoid an arbitrary splitting of the sample period; and second, since it does not require us to cluster the observations according to some pre-specified criterion, it does not confine attention to monetary policy but encompasses more general issues, such as credibility, changes in operating procedures and shifts in stance.

After having developed the VAR model with Markov-switching, we estimate the CAPM relation for five portfolios of Italian industry (manufacturing, services, banks, finance and insurance).[11] We then divide the β of each industry portfolio into the components related to the different state variables, following the methodology presented by Campbell and Mei (1993).

In this framework, risky assets provide a complete hedge against expected inflation if the following three conditions are satisfied: the nominal returns on short-term riskless rates move one-to-one with expected inflation; the β of a stock is not affected by anticipated changes in the price index; and the expected component of the excess return on the market portfolio is not correlated with expected inflation. These conditions also allow a test of the Fisher hypothesis, provided that it holds for the riskless asset.

The empirical framework can also deal with a more general assumption, i.e. that the Fisher hypothesis need not necessarily hold for the riskless asset. As will be shown in Section 3.2, the effect of expected inflation on nominal returns is estimated for every asset and the degree of coverage provided by stock returns could turn out to be different from that achieved on the short-term asset.

3.1 The Campbell and Shiller decomposition and the Markov-switching VAR

The model uses a log-linear approximation of the present value relation proposed by Campbell and Shiller. The basic equation links the unexpected stock excess return to changes in the rational expectation of future dividend growth, real interest rates, inflation and future excess returns. If e_{t+1} is the excess return on a stock held from the end of period t to the end of period t+1, d_{t+1} the log real dividend paid during period t+1, r_{t+1} the short-term riskless real interest rate and π_{t+1} the inflation rate, then the equation is:

$$e_{t+1} - E_t e_{t+1} = (E_{t+1} - E_t)\left\{ \sum_{j=0}^{\infty} \rho^j \Delta d_{t+1+j} - \sum_{j=0}^{\infty} \rho^j r_{t+1+j} - \sum_{j=0}^{\infty} \rho^j \pi_{t+1+j} - \sum_{j=1}^{\infty} \rho^j e_{t+1+j} \right\} \tag{3.1}$$

which can be also written in a more compact form as:

$$\tilde{e}_{i,t+1} = \tilde{e}_{di,t+1} - \tilde{e}_{r,t+1} - \tilde{e}_{\pi,t+1} - \tilde{e}_{ei,t+1} \tag{3.2}$$

Once the above asset return components have been computed, it is straightforward to derive the βs between innovations in stock excess returns and the state variables. This means that the

[11] Panetta and Zautzik (1990) show that the CAPM fits Italian stock market data quite well and that there is not much to gain in using a multi-factor model to explain excess returns on risky assets.

latter are used as factors, as in Chen, Ross and Roll (1976) and Fearson (1990). From (3.2), it follows that:

$$\beta_{i,m} \equiv \frac{Cov(\tilde{e}_{di,t},\tilde{e}_{m,t})}{Var(\tilde{e}_{m,t})} - \frac{Cov(\tilde{e}_{r,t},\tilde{e}_{m,t})}{Var(\tilde{e}_{m,t})} - \frac{Cov(\tilde{e}_{\pi i,t},\tilde{e}_{m,t})}{Var(\tilde{e}_{m,t})} - \frac{Cov(\tilde{e}_{ei,t},\tilde{e}_{m,t})}{Var(\tilde{e}_{m,t})} \equiv \beta_{di,m} - \beta_{r,m} - \beta_{\pi,m} - \beta_{ei,m} \quad (3.3)$$

If one is willing to assume that expectations of future returns are accurately described by the CAPM, then it is possible to substitute out the last term in (3.3). The decomposition of the overall β therefore becomes:

$$\beta_{i,m} = \frac{\beta_{di,m} - \beta_{r,m} - \beta_{\pi,m}}{1+\beta_{em,m}} \quad (3.4)$$

To become operational, the above formulae require a number of hypotheses about the mechanism driving expectations formation. The solution adopted by Campbell and Shiller is to assume that forecasts of excess returns can be approximated by a linear combination of a vector of state variables, x_t,[12] and that the law of motion of these variables can be adequately described by a VAR process.

We have made the further assumption that VAR coefficients are not constant throughout the sample period but rather are subject to occasional discrete shifts; the probability law governing these shifts is represented by a two-state Markov chain. In accordance with the literature surveyed above, we assume that only two regimes are allowed.

The state-space representation of the Markov-switching VAR is the following:

$$x_t = \Pi_{s_t} x_{t-1} + \tilde{x}_t \quad (3.5)$$

$$\xi_t = F\xi_{t-1} + \eta_t \quad (3.6)$$

where s_t is an unobserved random variable that takes the values 1 or 2 depending on in which regime the process is at time t; ξ_t is a two-element random vector, equal to $[1,0]$ if $s_t = 1$ and $[0,1]$ otherwise; $F \equiv \{p_{ij}\}_{i,j=1,2}$ is the transition matrix and p_{ij} is the probability that $s_t = j$ given that $s_{t-1} = i$. The assumption of a first-order centred VAR is not at all restrictive, since (3.5) has to be interpreted as the companion form representation of the process.[13]

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[12] The VAR approximation of the mechanism of expectations formation faces at least two problems: first, expectations concern variables which are realised only over long periods of time; second, investors may have information which is not available to the econometrician or cannot be summarised by means of aggregated variables. The first problem can be handled by using the VAR expressions for multi-period forecasts, while the second does create difficulties. The only case where investors' superior information does not distort the analysis occurs when only one component of an asset price is time varying, since then the asset price itself contains all the relevant information about that component. In the general case, the VAR results must be interpreted cautiously, conditional on whatever information is included in the system.

[13] A tricky issue concerned with the analysis of the βs is the proper evaluation of the precision of the estimate. The approach suggested by Campbell is to treat the VAR coefficients and the elements of the covariance matrix of the residuals as parameters to be jointly estimated by GMM. The GMM parameter estimates are numerically identical to OLS ones, but GMM delivers a heteroskedastic-consistent covariance matrix V for the entire set of coefficients. Since the βs can be thought of as non-linear functions $f(\gamma)$ of the vector γ of parameters of the model, their variances turn out to be a quadratic form in the first derivative of $f(\gamma)$ and V. In the present set-up, this procedure is clearly unfeasible: the vector of

The use of a Markov process turns out to be useful on several grounds:

- As stressed forcefully by Sims (1982) and Cooley et al. (1984), it is at least doubtful whether changes in the policy framework should be characterised as permanent changes in the parameters of a reaction function, since genuine changes in regime are rare events. From past experience, economic agents know the menu of choices available to the policymakers and form expectations accordingly, taking into account all the possible outcomes. In other words, they have a probability distribution ranging over all possible policy rules and use it to forecast the behaviour of policymakers.
- A Markov-switching model is flexible enough to encompass once-and-for-all structural changes as well as period-by-period time-varying models. The first case corresponds to each state being a so-called absorbing state, which lasts forever once reached; the second can be approximated by assuming that there exist a large number of states. Any intermediate case can be obtained by appropriately choosing the parameters of the transition matrix.
- Finally, relying on a statistical procedure to split the sample period avoids arbitrary and unnecessarily restrictive assumptions. Monetary policy, which in the literature is usually considered responsible for regime shifts, may not be the only source of instability. Fiscal as well as incomes policies may play a similar role, not to mention the effects deriving from changes in the institutional framework within which economic agents operate. Focusing attention on only one source of instability may be unduly restrictive and could strongly bias the results. Using a statistical technique such as a Markov-switching model has the advantage of allowing the data to speak for themselves; furthermore, the interpretation of the odds attributed to a given regime in each time period provides a genuine test of the reliability of the method.

3.2 The results

The simple tests presented in Section 2 do not support the hypothesis that nominal yields on short-term government securities fully incorporate agents' inflation forecasts (Tables 1a and 1b); at a 5% confidence level, the Fisher hypothesis (together with market efficiency and the null of no correlation between expected inflation and the real rate) is almost always rejected.[14]

However, short-term assets provide partial insurance against expected inflation: the estimated effect on monthly data ranges between 0.41 and 0.64 and does not change significantly either with the maturity of the short-term asset or with the frequency of the data. However, these values are not stable throughout the estimation period.

The splitting of the sample period provided by the Markov-switching algorithm is shown in Figure 2. It is apparent from the graph that the second regime becomes the dominant one in the last quarter of 1988, after a two-year transition period. The interpretation of the change in regime can be clearly related to policy actions and changes in the institutional framework:

- after the realignment of the lira in January 1987, the exchange rate commitment became more credible and no other changes in the central parity of the Italian currency took place until the

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estimated coefficients has more than 150 elements and the matrix of first derivatives has more than 20,000, not to mention the fact that it is not at all easy to find and differentiate the function relating the βs to the VAR parameters. The solution adopted in this paper is to consider the problem as a special case of the general issue of efficiently and consistently estimating second moments in a model with generated regressors (McKenzie and McAleer (1990) and Pagan (1984)). It is well known that the application of OLS to models with generated regressors will generally be inefficient and lead to inconsistent estimates of the standard errors of the regressor coefficients. A convenient way out of this problem, which has been adopted in this paper, is to allow for non-spherical errors and to use a GLS-type estimator; a simpler alternative is to compute the t-statistics by using a consistent estimate of the covariance matrix of the error term.

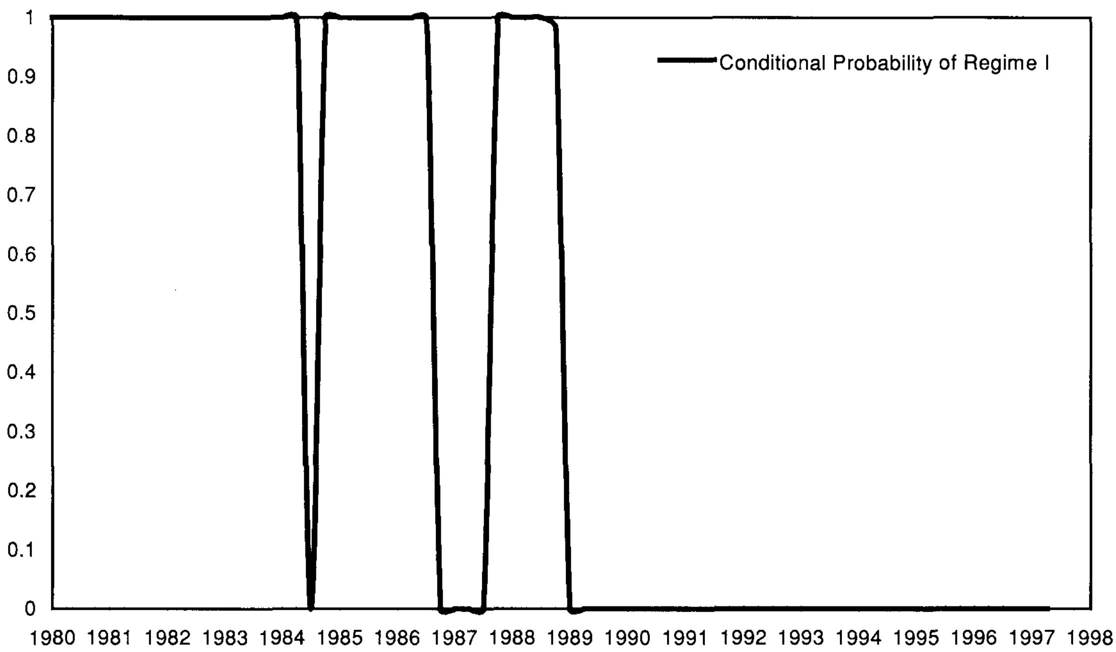
14 This result is robust to different measures of expected inflation and holds for both monthly and quarterly data.

exit of the lira from the Exchange Rate Mechanism of the EMS in September 1992. Between 1987 and 1990, capital movements were progressively liberalised to comply with the requirements set by the EC for the Single Market. The most important measure became effective on October 1988, when all capital movements, except those involving monetary instruments, were liberalised (see Passacantando (1996)). In January 1990, the fluctuation band was narrowed from 6 to 2.25% and the remaining capital controls were completely abolished by April of the same year;

- in May 1988, a screen-based secondary market for government securities was introduced. Between July of that year and March 1989, the floor price for Treasury bills in the primary market was abolished for all maturities. In February 1990, a screen-based market for interbank deposits was launched. In October, banks were allowed to mobilise part of their compulsory reserves. All of these reforms contributed to shifting the conduct of monetary policy from administrative controls to market-oriented procedures.
- incomes policy can also be a factor in a regime change. In the first half of 1984, wage increases were agreed on the basis of a planned rate of inflation rather than relying on a backward-looking indexation mechanism. In 1986, the wage indexation mechanism was further modified by reducing the overall degree of coverage and lowering the frequency of the adjustment (from 3 to 6 months).

Figure 2

Conditional probability of the Italian economy being in regime I



The smoothed probabilities associated with the two regimes also indicate that a reversal of the first occurred in the period 1987:III-1988:III. This shift coincides to a large extent with the reintroduction of controls on bank lending (from September 1987 to March 1988).

The estimates of the VAR model with Markov-switching are reported in Table 2. The effects of past values of inflation on the stock excess return provide a measure of the relationship between expected inflation and the premium requested on stocks. The coefficients indicate that past inflation does not contribute to explaining movements in the overall risk premium; neither in the first nor in the second regime does lagged inflation seem to affect the current excess return on the market portfolio.

Table 2

Double regime Markov-switching VAR model

Equation for market excess returns

	Regime I	Regime II
constant	-0.3599	-0.0836
	(0.1026)	(0.0627)
e_{t-1}	-0.0982	0.2178
	(0.1121)	(0.2123)
r_{t-1}	-9.5344	-1.2027
	(3.8742)	(6.2208)
dy_{t-1}	0.3949	27.1170
	(48.9280)	(55.8980)
π_{t-1}	2.7243	1.4835
	(1.9078)	(2.4867)
Δip_{t-1}	-2.0274	-1.1850
	(0.5814)	(0.3925)
e_{t-2}	0.3669	0.2252
	(0.1142)	(0.1264)
r_{t-2}	6.5005	7.2364
	(3.4563)	(6.5171)
dy_{t-2}	66.7410	-20.8520
	(16.635)	(35.8470)
π_{t-2}	-0.7223	-4.7145
	(10.6239)	(2.2791)
Δip_{t-2}	-0.3231	0.6242
	(0.8121)	(0.4036)

Equation for the short-term real riskless rate

	Regime I	Regime II
constant	0.01246	0.0032
	(0.0028)	(0.0022)
e_{t-1}	-0.0003	-0.0068
	(0.0028)	(0.0063)
r_{t-1}	0.2084	0.7772
	(0.1670)	(0.1640)
dy_{t-1}	-0.7231	0.0560
	(0.6479)	(1.2973)
π_{t-1}	-0.0047	-0.0484
	(0.0479)	(0.0687)
Δip_{t-1}	-0.0032	0.0183
	(0.0221)	(0.0161)
e_{t-2}	-0.0017	-0.0022
	(0.0025)	(0.0043)
r_{t-2}	0.3572	-0.2744
	(0.1024)	(0.1528)
dy_{t-2}	-0.2144	0.3691
	(0.6021)	(1.1270)
π_{t-2}	-0.0536	0.1040
	(0.0412)	(0.0722)
Δip_{t-2}	0.0533	-0.0130
	(0.0208)	(0.0122)

Equation for the dividend yield

	Regime I	Regime II
constant	0.0023	0.0007
	(0.0007)	(0.0004)
e_{t-1}	-0.0034	-0.0036
	(0.0014)	(0.0017)
r_{t-1}	-0.0099	-0.0281
	(0.0545)	(0.0423)
dy_{t-1}	0.7999	0.7405
	(0.2454)	(0.2965)
π_{t-1}	-0.0362	0.0090
	(0.0242)	(0.0129)
Δip_{t-1}	0.0152	0.0052
	(0.0059)	(0.0036)
e_{t-2}	-0.0004	-0.0011
	(0.0010)	(0.0012)
r_{t-2}	0.0060	-0.0156
	(0.0409)	(0.0325)
dy_{t-2}	-0.1163	0.1116
	(0.2649)	(0.2559)
π_{t-2}	0.0189	0.0474
	(0.0143)	(0.0126)
Δip_{t-2}	-0.0072	-0.0045
	(0.0056)	(0.0027)

Equation for the index of industrial production

	Regime I	Regime II
Constant	0.0401	0.0620
	(0.0271)	(0.0278)
e_{t-1}	-0.0341	0.0943
	(0.0292)	(0.0704)
r_{t-1}	-0.6866	3.9414
	(0.9705)	(1.8218)
dy_{t-1}	-4.5008	-18.7210
	(3.8494)	(12.6160)
π_{t-1}	0.9620	-0.9156
	(0.3391)	(0.6646)
Δip_{t-1}	0.0959	-0.2467
	(0.1354)	(0.1561)
e_{t-2}	0.0508	-0.0534
	(0.0262)	(0.0520)
r_{t-2}	0.1406	-6.0491
	(0.5458)	(1.8904)
dy_{t-2}	2.9878	13.8580
	(3.1308)	(12.0130)
π_{t-2}	-1.7657	1.0131
	(0.2922)	(0.6461)
Δip_{t-2}	0.0745	-0.2363
	(0.1517)	(0.1710)

Equation for the rate of inflation

	Regime I	Regime II
constant	0.0068	0.0081
	(0.0065)	(0.0032)
e_{t-1}	0.0025	-0.0047
	(0.0108)	(0.0105)
r_{t-1}	0.2634	-0.5306
	(0.359)	(.3026)
dy_{t-1}	4.6364	1.6690
	(1.8998)	(2.0778)
π_{t-1}	0.3537	0.1591
	(0.1944)	(0.0971)
Δip_{t-1}	-0.2553	0.0539
	(0.0378)	(0.0215)
e_{t-2}	0.0061	-0.0118
	(0.0104)	(0.0057)
r_{t-2}	-0.7168	0.0644
	(0.3204)	(0.2291)
dy_{t-2}	-3.8665	-0.4569
	(1.8501)	(1.9652)
π_{t-2}	0.3333	0.0482
	(0.1615)	(0.1012)
Δip_{t-2}	-0.1055	0.0425
	(0.0751)	(0.02164)

Notes: The VAR model is estimated on quarterly data, for the period 1979:4-1997:1. Numbers in parentheses are coefficient standard errors, calculated according to the formulas suggested in Hamilton (1996). The variables are defined as follows: e_t is the excess return on the market portfolio, r_t the riskless short-term rate, dy_t the dividend yield, π_t is the rate of inflation, and Δip_t is the first difference of the logarithm of the index of industrial production.

Table 3

Italian stock exchange sub-indexes:
Campbell and Shiller's decomposition of the βs with respect to the market portfolio

	Manufacturing		Services		Credit		Finance		Insurance	
	Regime I	Regime II	Regime I	Regime II	Regime I	Regime II	Regime I	Regime II	Regime I	Regime II
β dividends	1.6392	2.2592	1.6215	2.0408	1.6808	1.9176	1.8999	2.1656	1.9771	1.7804
	(15.859)	(48.784)	(19.127)	(20.718)	(7.877)	(16.386)	(20.795)	(42.875)	(18.247)	(20.464)
β real rate	0.89145	1.0185	0.89059	1.0185	0.88596	1.0184	0.88437	1.0183	0.87843	1.0180
	(27.460)	(40.603)	(27.341)	(40.612)	(26.718)	(40.494)	(26.512)	(40.241)	(25.765)	(39.284)
β inflation	-0.00774	-0.01091	-0.00776	-0.01091	-0.00783	-0.01093	-0.00786	-0.01097	-0.00795	-0.01112
	(-5.426)	(-18.233)	(-5.426)	(-18.232)	(-5.427)	(-18.240)	(-5.427)	(-18.258)	(-5.429)	(-18.314)
β future excess returns	-0.11964	0.06833	-0.17944	0.04022	-0.09882	0.00822	-0.13035	0.09012	-0.08024	-0.02533
	(-4.455)	(3.361)	(-6.580)	(-0.983)	(-2.210)	(-0.464)	(-3.218)	(5.328)	(-2.108)	(-0.593)
β total	0.87516	1.1833	0.91808	1.0735	0.9015	0.9183	1.1538	1.0681	1.1869	0.79889
	(7.675)	(22.629)	(18.481)	(13.055)	(5.407)	(7.819)	(19.350)	(17.819)	(16.374)	(13.769)

Notes: According to the present value relation, the β of a sub-index with respect to the market portfolio can be decomposed as follows (see equation (3.3)):

$\beta_{total} = \beta_{dividend} - \beta_{real\ rate} - \beta_{inflation} - \beta_{future\ excess\ return}$. Numbers in parentheses are t-statistics.

The two VAR models provide two sets of residuals, which have been used to compute two sets of βs for five industry portfolios, one for each regime (see the last row in Table 3); each β has then been identified as the combination of four components: the real rate, inflation, dividends and excess returns (see equation (3.3)).

Two general remarks on the interpretation of the βs are in order. Since the five portfolios sum up to the whole market and the β for the market is one, the overall βs, shown in the last row of the table, increase in some cases and decrease in others when moving from one regime to another. A further warning is due: since the VAR is not identified, innovations in the state variables are not orthogonalised and the *ceteris paribus* clause cannot be applied in interpreting the βs. This means that the residuals of the VAR equations do not identify exogenous, idiosyncratic shocks to the state variables, but rather represent the unexpected components in the state variables with respect to the previous period information set.

The share of the β of a portfolio attributed to news about the future short-term real interest rate measures the main channel whereby monetary policy affects stock prices, while the β related to inflation provides a quantitative assessment of the effect of inflation innovations on stock excess returns: if the latter did not exert a significant influence on stock excess returns, the value of β_π would be very low and barely significant.

In most cases, all the βs show the same sign across portfolios and the same ranking (in absolute value) across regimes. Compared with the results in Campbell and Mei, all βs show the expected sign: positive for cash-flow and real rate and negative for inflation and future excess return.[15] The differences in the βs across regimes are substantial and statistically significant,[16] showing the existence of tight links between policy actions and market behaviour and supporting the sample splitting induced as a Markovian latent variable.

The dividend component is positive and by far much larger than the other components. In Campbell and Mei, by contrast, the cash-flow βs are always smaller than those related to future excess returns. The size of the dividend component may be overstated, because it is computed as a residual; indeed, one might suspect that the harder portfolio returns are to forecast, the more important the dividend component becomes. But this cannot be the whole story for at least two reasons. First, as is observed by Cambell and Mei, there is no incontrovertible evidence that the fit of the regressions for portfolios' excess returns, as measured by the adjusted multiple correlation coefficient, is negatively related to the size of the residual dividend component. Second, even if $\tilde{e}_{di,t+1}$ is large, there is no guarantee whatsoever that β_d is also large, since most of the variation in cash-flows could be idiosyncratic.

It is worth stressing that β_d changes dramatically between the two sub-samples. The significant increase in the second regime may reflect factors peculiar to the Italian market. Until the mid-eighties companies mostly raised funds by borrowing from banks, thanks to a cheap credit; only rarely did bond or equity issues represent a significant source of financing. In the nineties, owing in part to higher real interest rates and banks' restructuring, an increasing number of companies turned to the international capital markets and thus had an incentive to pursue a dividend policy more akin to those in countries with more developed stock markets. As a result, dividends themselves have become a binding constraint for companies, influencing their investment projects. Another event may have strengthened this process. Starting in the late eighties, small and medium-sized firms have been listed on the Milan stock exchange. Because their capacity to borrow in international capital markets is

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[15] However, only in few cases is the latter positive.

[16] Two different tests have been computed: in the first case, it has been assumed that residuals are homoskedastic while, in the second, time-varying second moments have been allowed.

limited, they have been forced to pay a great deal of attention to dividend policy. At the same time, two other developments may have altered the sensitivity of stock returns to cash-flow news: the introduction of new financial intermediaries, namely mutual funds, and changes in the tax code.

The estimated values for the βs associated with the real interest rate are positive and increase between the first and the second regime;[17] but the magnitude is greater than that computed on US data. The sign of the correlation is as expected for two reasons. First, since stock prices are forward-looking, they can react to information that is used by the central bank for the conduct of monetary policy. Second, changes in asset prices may have a direct impact on aggregate demand, via both consumption and investment expenditure; whenever the central bank is not confident that this appreciation is fully justified by changes in fundamentals, it may choose to intervene to avoid excessive price and output variability. The increase in β_r in the second regime reinforces this interpretation, as the second half of the sample is characterised by a more restrictive monetary policy stance; increased sensitivity of stock returns to real rates in a tighter monetary environment is also one of the main implications of Söderlind's model.

The estimated effects of news about inflation are negative. Although their size is of second-order compared with the cash-flow and real rate components, they are all highly significant, showing that unexpected inflation exerts some influence on the excess return required on stock portfolios. This evidence is consistent with the Mundell-Tobin effect: upward revisions of agents' inflation forecasts result in a rebalancing of portfolios from money to other assets. In moving from the first to the second regime, the variance of inflation innovations decreases (as is to be expected when monetary policy assigns more weight to inflation targets), while β_π increases. This finding can be related to the greater openness of the Italian economy in the nineties, which has increased the costs of inflationary shocks for most of Italy's listed companies.

The β component associated with future excess returns is generally small and in most cases not statistically significant. This contrasts with the evidence presented by Campbell and Mei for US data, in which, on average, this component is the largest. A possible explanation may be the weak persistence of Italian stock returns, which stands in stark contrast to the US data.

These results provide a first clue about the influence of inflation on stock returns. However, the assumption that portfolio sensitivity to systematic risk is constant within each regime and not allowed to respond to changes in inflation may be unwarranted. This may introduce a bias in the measure of the Fisher effect. In order to test time variation in the overall βs, we have replicated the analysis of Ferson and Schadt (1996) by regressing the innovation in each portfolio's excess returns on the innovation on the markets excess return and the cross products of the latter with each element of the Campbell and Shiller decomposition.[18] The results of the estimates tend to reject time variation in the βs, thus providing additional support to the previous findings.

All in all, the evidence supports the claim that in the last twenty years Italian stocks have not provided a better hedge against inflation than government securities, even when the effects of policy actions on market expectations are taken into account.

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[17] The sensitivity of the market return to inflation and real interest rate news is approximately the same for all portfolios: marginal differences are due to the discount factor, which is related to the yield ratio. As equation (A.11) in Appendix 2 points out, only the parameter ρ is different across portfolios.

[18] All variables, except the innovation on the excess return on the market portfolio, have been lagged once, in order to ensure that they represent commonly available information.

Summary and conclusions

This paper builds on two main ideas.

(i) Testing the Fisher hypothesis by simple projections of nominal returns on expected inflation is misleading, since those regressions are reduced-form models, powerfully affected by changes in both policy actions and in the institutional framework. The sign and size of the parameter associated with expected inflation can take any value, depending on which variables are added among the regressors. Moreover, the menu of omitted variables is endless, since, in principle, any variable appearing in a structural macro model can be relevant to changes in nominal returns. A structural model is therefore the proper framework within which to analyse the correlation between returns and inflation;

(ii) As is clearly pointed out in the literature on inflation and stock returns, monetary policy must be dealt with to provide a proper account of the relevance of the Fisher hypothesis to the stock market. However, given that the potential sources of instability in the relation between asset returns and inflation are not limited to monetary policy but also include fiscal and income policies as well as changes in the institutional environment, imposing the splitting of the sample among different regimes on the grounds of a priori evaluations would not appear to be the safest and most valuable modelling strategy. The alternative proposed in this paper is to model regime shifts as a stochastic latent variable, with non-sample information not used in setting up the model, but rather in interpreting the results. The advantage is that while no information is discarded, results are not biased by untested assumptions and due attention can be paid not just to monetary policy but to other policy factors.

As a first step in the empirical analysis, we run simple tests of the Fisher hypothesis for Italian Treasury bills and bonds on a sample covering the last twenty years. We apply the same test to equities, to check whether the negative relationships between inflation and stock yields found by Fama and Schwert for US data applies to Italian data as well. In line with evidence for other countries, we find that government securities provide only a partial hedge against expected inflation, while the estimated relationship for common stocks proves inconclusive, due to instability in the parameters.

Within the framework of the CAPM and the log-linear present-value model suggested by Campbell and Shiller, we then test for the influence of inflation on the excess returns required by investors in equities, with a separate analysis of portfolio βs and factor risk. To generate innovations in the state variables, we assume that financial markets form expectations about the relevant macroeconomic variables by means of a VAR model and that the parameters of the expectation formation mechanism change across policy regimes. Finally, we compute βs for a number of industry portfolios and use this decomposition to make inferences about policy actions and the Fisher effect.

The main conclusions of this section are the following:

- sample evidence indicates a shift in the policy environment in the second half of the eighties, when the exchange rate commitment became more binding, monetary control was definitively based on market instruments, and incomes policies became stricter;
- the evidence on short-term assets does not support the Fisher hypothesis, but expected inflation is widely incorporated in short-term interest rates;
- there is no evidence of an influence of inflation forecasts on the market excess return, though the evidence is less clear-cut for the second regime;
- inflation does not seem to have a significant influence on stock excess returns and no time variation in the βs induced by movements in inflation was detected.

Once all the channels through which inflation affects stock returns are taken into account, it turns out that in the last two decades stocks have not significantly outperformed government securities as hedges against inflation.

Appendix 1: Data description

Industrial production: the index of industrial production refers to manufacturing, marketable services and energy. It is collected monthly and adjusted for the number of working days; seasonal adjustment is by an X11-Arima filter.

Inflation: log of the first difference of the cost of living index, net of tobacco products. The index uses a basket of 290 items, which refer to more than five hundred goods and services; data are collected monthly in the capitals of the twenty Italian regions.

Forum-ME survey of inflation expectations: since 1952, the Italian magazine *Mondo Economico* (ME) has conducted surveys on expectations of inflation. The respondents are selected within four main categories of economic agents: managers and executives in industrial, financial and commercial sectors, and business economists. The surveys are conducted by means of an anonymous mail questionnaire and answers have to fall into one of a number of pre-selected intervals (the lower and upper ones being of course open intervals). Until 1981, the survey was run twice a year and covered a six-month forecasting horizon; thereafter, it has been quarterly, with a corresponding shortening of the time frame.

Treasury bills: allotment rate at end-of-month auctions gross of withholding tax. Until June 1981, the Bank of Italy was committed to act as residual buyer for unsubscribed bills. Competitive-bid auctions replaced uniform price auctions in May 1983 for 3-month bills, in May 1984 for 6-month bills and in February 1988 for 12-month bills; for competitive-bid auctions, the yield is the weighted average allotment rate. A floor price for each auction was fixed by the Treasury until June 1988 for 3-month bills and February 1989 for 6 and 12-month bills.

Treasury bonds: average yield of the BTPs with at least one year to maturity traded on the Italian stock exchange, gross of withholding tax.

Dividend yield: total dividends paid over the previous year relative to the current stock price; the latter is computed on the basis of end-of-month closing prices. Data refer to shares of Italian companies listed on the Italian stock exchange.

Stock returns: holding period returns computed on the basis of value-weighted portfolio indexes; Italian listed companies.

Appendix 2: An approximate present-value model with a stochastic discount factor

The model suggested by Campbell and Shiller is a modified version of the present value equation in real terms, relating unexpected returns to changing expectations of future cash flows, real interest rates and excess return.[19] Since the model is derived from a dynamic accounting identity, it is not conditional on any particular asset pricing model; but if one is willing to impose a theoretical structure, it is possible to cancel future required returns and to relate unexpected excess returns to future cash flows and real interest rates only.

The model is derived from the Gordon present value relation, by disposing of the assumption concerning the constancy of the discount factor. Though the relaxation of this hypothesis improves the accuracy of the model, it creates problems of its own, since time-variability of stock returns introduces non-linearities. To overcome this, Campbell and Shiller propose taking logs and linearising the present value relation. The approximate equation is then solved forward, imposing a "no rational bubble" terminal condition.

Starting from the definition of gross stock returns and taking logs, we have:

$$\begin{aligned}\log(1+H_{t+1}) &\equiv \log(P_{t+1}+D_{t+1})-\log(P_t)=\log(P_{t+1})+\log\left(1+\frac{D_{t+1}}{P_{t+1}}\right)-\log(P_t)\\ &= p_{t+1}-p_t+\log(1+\exp(d_{t+1}-p_{t+1})) \qquad \text{(A.1)}\end{aligned}$$

where H_{t+1}, D_{t+1} and P_{t+1} are, respectively, the real return, the dividend and the price of the stock or portfolio we are considering (by the standard convention, logs of variables are denoted with lower-case letters). The last term on the right-hand side is a nonlinear function of the log dividend-price ratio, which can be approximated around the mean using a first-order Taylor expansion:

$$\begin{aligned}\log(1+\exp(d_{t+1}-p_{t+1})) &\approx \log(1+\exp(\bar{d}-\bar{p}))+\frac{\exp(\bar{d}-\bar{p})}{1+\exp(\bar{d}-\bar{p})}\left[(d_{t+1}-p_{t+1})-(\bar{d}-\bar{p})\right]\\ &= -\log(\rho)-(1-\rho)\log\left(\frac{1}{\rho}-1\right)+\rho p_{t+1}+(1-\rho)d_{t+1}-p_t \qquad \text{(A.2)}\end{aligned}$$

where $\rho \equiv \frac{1}{1+\exp(\bar{d}-\bar{p})} \equiv \frac{\bar{P}}{\bar{P}+\bar{D}}$ (the bar indicates sample means). ρ is a number close to 1 and plays the role of a weighting factor. The reason is intuitive: the dividend is much smaller than the stock price, so a given percentage change in the dividend component must have a much smaller effect than the same variation in the price. Substituting (A.2) into (A.1) and solving forward yields:

$$p_t = -\frac{\log(\rho)}{1-\rho}-\log\left(\frac{1}{\rho}-1\right)+\sum_{j=0}^{\infty}\rho^j\left[(1-\rho)d_{t+1+j}-h_{t+1+j}\right] \qquad \text{(A.3)}$$

where the definition $\log(1+H_{t+1}) \equiv h_{t+1}$ has been used. This equation is to be interpreted as a dynamic accounting relation, obtained by approximating an identity; it holds *ex post* but also *ex ante*, once future realisations of dividends and returns are replaced by their expected values:

[19] A thorough treatment of the present-value relation can be found in Chapter 7 of Campbell et al. (1997).

$$p_t = -\frac{\log(\rho)}{1-\rho} - \log\left(\frac{1}{\rho}-1\right) + E_t \sum_{j=0}^{\infty} \rho^j \left[(1-\rho)d_{t+1+j} - h_{t+1+j}\right] \tag{A.4}$$

Rearranging (A.1) so that the rate of return is the left-hand variable and substituting (A.4) for both p_t and p_{t+1}, we can write asset returns as linear combinations of revisions in expectations:

$$h_{t+1} - E_t h_{t+1} = (E_{t+1} - E_t)\left\{ \sum_{j=0}^{\infty} \rho^j \Delta d_{t+1+j} - \sum_{j=1}^{\infty} \rho^j h_{t+1+j} \right\} \tag{A.5}$$

This equation links the unexpected real stock return in period t+1 to changes in the rational expectation of future dividend growth and future stock returns. Equation (A.5) must be interpreted as a consistency condition for expectations; it states that if the unexpected stock return is negative, then either expected future dividend growth must be lower or expected future stock returns must be higher, or both. The discount factor ρ indicates that the further in the future the expectation of a change in returns is, the smaller is the change in today's stock price.

For many purposes it is convenient to work with excess stock returns. If the log real interest rate on a riskless short-term security is r_{t+1}, then the excess return is just $e_{t+1} \equiv h_{t+1} - r_{t+1}$. Substituting this expression into (A.5) provides the following consistency condition:

$$e_{t+1} - E_t e_{t+1} = (E_{t+1} - E_t)\sum_{j=0}^{\infty} \rho^j \Delta d_{t+1+j} - (E_{t+1} - E_t)\sum_{j=0}^{\infty} \rho^j r_{t+1+j} - (E_{t+1} - E_t)\sum_{j=1}^{\infty} \rho^j e_{t+1+j} \tag{A.6}$$

In this paper, we have used a slightly modified version of this equation, obtained by taking a present value relation expressed in nominal rather than real terms as a starting point. In this case, (A.6) becomes:

$$e_{t+1} - E_t e_{t+1} = (E_{t+1} - E_t)\left\{ \sum_{j=0}^{\infty} \rho^j \Delta d_{t+1+j} - \sum_{j=0}^{\infty} \rho^j r_{t+1+j} - \sum_{j=0}^{\infty} \rho^j \pi_{t+1+j} - \sum_{j=1}^{\infty} \rho^j e_{t+1+j} \right\} \tag{A.7}$$

which can also be written in a more compact form as:

$$\tilde{e}_{i,t+1} = \tilde{e}_{di,t+1} - \tilde{e}_{r,t+1} - \tilde{e}_{\pi,t+1} - \tilde{e}_{ei,t+1} \tag{A.8}$$

(The meaning of these terms is evident, by comparing (A.8) with (A.7).)

The excess return on a portfolio is assumed to be predictable by means of a projection on a vector of state variables x_t:

$$e_{i,t+1} = a_i' x_t + \tilde{e}_{t+1} \tag{A.9}$$

where a_i' is a vector of projection coefficients and $\tilde{e}_{t+1}$ is the unexpected component of the excess return.

To become operational, the above formulas require some hypotheses concerning the mechanism that drives expectation formation. The solution adopted by Campbell and Shiller is to assume that the law of motion of the state variables can be adequately described by a VAR process:

$$x_{t+1} = \Pi x_t + \tilde{x}_{t+1} \tag{A.10}$$

where $\tilde{x}_{t+1}$ is the innovation in the state vector. To allow for higher order processes or deterministic components, one must suitably augment the dimension of the vector of state variables. The first three elements of x_t are the excess return on the market, the real return on a short-term Treasury bill and the rate of inflation; the other components are selected from variables that are known to the market by time t and that have been shown in the literature to have some explanatory power for future returns:[20] for example, the dividend yield, the slope of the term structure and the default spread. Given the VAR model, revisions in rational expectations of the state variables are provided by the expression:

$$(E_{t+1} - E_t)x_{t+1+j} = \Pi^j \tilde{x}_{t+1} \tag{A.11}$$

Equation (A.11) enables us to compute the right-hand terms in (A.6) and (A.7). If i_j indicates the vector that picks the j-th component of $\tilde{x}_{t+1}$, the following equations hold:

$$\tilde{e}_{em,t+1} = \rho a_1'(I - \rho\Pi)^{-1}\tilde{x}_{t+1}$$

$$\tilde{e}_{ei,t+1} = \rho a_i'(I - \rho\Pi)^{-1}\tilde{x}_{t+1}$$

$$\tilde{e}_{r,t+1} = i_2'(I - \rho\Pi)^{-1}\tilde{x}_{t+1}$$

$$\tilde{e}_{\pi,t+1} = i_3'(I - \rho\Pi)^{-1}\tilde{x}_{t+1}$$

$$\tilde{e}_{di,t+1} = \tilde{e}_{i,t+1} + (i_2' + \rho a_i')(I - \rho\Pi)^{-1}\tilde{x}_{t+1} \tag{A.12}$$

The component associated with innovation in the path of dividend growth is computed as a residual and is therefore likely to be overstated. However, the sign of the bias is uncertain, since it will depend on the covariances between omitted and included variables.

Once the above asset return components have been computed, it is straightforward to derive the βs between innovations in stock excess returns and in the state variables. This means that the latter are used as factors, as in Chen, Ross and Roll (1976) and Fearson (1990). From (A.7), it follows that:

$$\beta_{i,m} \equiv \frac{Cov(\tilde{e}_{di,t},\tilde{e}_{m,t})}{Var(\tilde{e}_{m,t})} - \frac{Cov(\tilde{e}_{r,t},\tilde{e}_{m,t})}{Var(\tilde{e}_{m,t})} - \frac{Cov(\tilde{e}_{\pi i,t},\tilde{e}_{m,t})}{Var(\tilde{e}_{m,t})} - \frac{Cov(\tilde{e}_{ei,t},\tilde{e}_{m,t})}{Var(\tilde{e}_{m,t})} \equiv \beta_{di,m} - \beta_{r,m} - \beta_{\pi,m} - \beta_{ei,m} \tag{A.13}$$

If one is willing to assume that expectations of future returns are well described by a simple CAPM, then the last term in (A.13) can be substituted out. The decomposition of the overall β thus becomes:

$$\beta_{i,m} = \frac{\beta_{di,m} - \beta_{r,m} - \beta_{\pi,m}}{1 + \beta_{em,m}} \tag{A.14}$$

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[20] See for instance Fama (1988), Fama and French (1988) and Boldrin et al. (1995).

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Japanese share prices

Shuichi Uemura and Takeshi Kimura

Preface

The bubble of the late eighties burst in the early nineties, plunging Japanese share prices into a prolonged slump that is in stark contrast to the rising share prices seen in other industrialised countries (Figure 1). This paper verifies, in light of conditions in the Japanese stock market, the role played by the information value of share prices, describes the distinguishing features of share price formation in Japan and makes some observations about the most recent share price slump. Below, the major points are summarised.

1) We begin by using Granger causality tests and time series correlations to verify the relationship between share prices and major economic indicators, finding that share prices lead several real economic indicators, including real GDP. We also use an econometric technique, called the Probit method, to verify the potential for share prices to forecast an economic recession, finding a certain degree of usefulness.

2) We next examine Japanese share price formation in the past, noting that a moving average of the rate of share price change evinces almost exactly the same trends as the rate of land price change. This indicates that there is a close relationship between share prices and land prices. Share price levels (market capitalisation) have been consistent with corporate net asset values when calculated in terms of reacquisition costs, and this trend held true even during the bubble period of the late eighties. Rising land prices made a considerable contribution to the increase in corporate net asset values during the late eighties, and it is likely that the unrealised profits on land, which contained a bubble, were translated directly into share price formation. This is consistent with the phenomenon seen in the nineties, when share prices have been slumping as land prices dropped.

3) Additionally, we use the "dividend discount model", one of the leading models for asset price determination, as a framework to consider the factors behind the recent share price slump. In the nineties, the difference between long-term interest rates and the earnings yield[1] – in other words, the yield spread – has continued to decline. This is basically a reflection of the decline in the expected growth rate of nominal earnings, but the expansion in the risk premium has also played a part. We regressed risk premium changes with several explanatory variables and found that the movements in the risk premium during the nineties can, for the most part, be explained by an expansion in credit risk. What is more, it is likely that falling land prices are behind this expansion in credit risk. Note that in recent years there has been a contrasting development in share prices between sectors that are respectively less and more vulnerable to land price drops.

4) It appears that the basic factors behind the slump in Japanese share prices are lower expected nominal growth rates and higher credit risks. Fundamentally, therefore, they are the after-effects of the land bubble. During this period we have also witnessed signs of structural changes in the stock market in the form of a less significant role being played by personal investors, a greater role of foreign investors, an unwinding of share crossholding relationships, and new emphasis on return on equity as an investment yardstick. It is not clear what influence these developments have had on share prices nor is the pace of change expected to accelerate in

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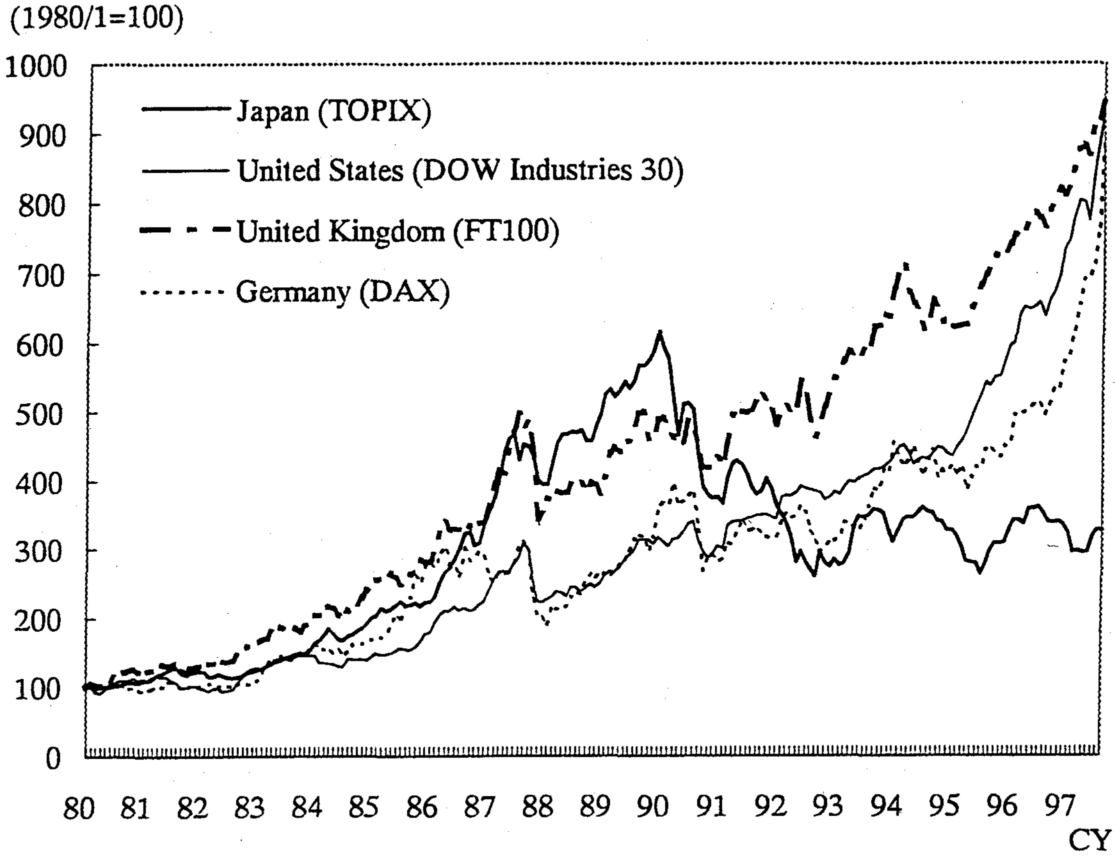
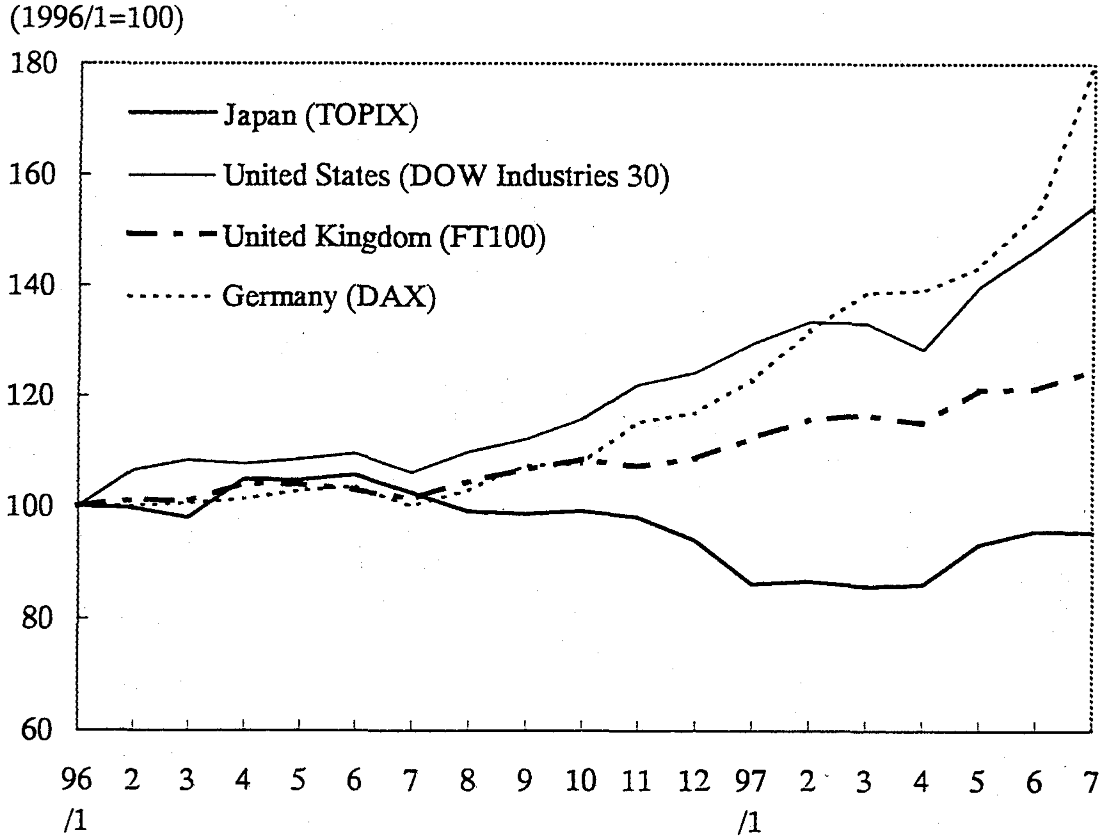
[1] The inverse of the price/earnings ratio; note that this paper uses a price/earnings ratio adjusted for cyclical factors and share crossholding relationships.

Figure 1

Share prices in industrialised countries





the future. However, policy makers who are looking at share prices will need to be aware of the influence that structural changes in the market may have on share price formation.

5) During the past year, share price movements have been unstable. The chief causes of this have been greater uncertainty about the economic future caused by fiscal consolidation and an expansion in credit risks as triggered by the after-effects of the bubble in the form of several corporate bankruptcies. The low yield spread would indicate that there is little room to consider Japanese shares as over-valued at current levels, but that does not mean the uncertainties over share prices will be resolved any time soon. This paper concludes that for Japanese share prices to recover in the future, four things will be required: recovery of the expected macroeconomic growth rate; relief from the high credit risks brought by falling land prices; more emphasis on shareholder values, such as the revision of dividend policies and improvement of return on equity (for example, by buying back shares from the market); and enhancements to market infrastructure, for example, better accounting and disclosure standards.

1. Share prices as an information variable

In this section, we use a number of statistical techniques to verify whether share price movements in Japan contain information regarding future economic conditions to a significant degree.

Granger test

We began by testing for Granger causality[2] using a two-variable VAR for the period from the first quarter 1970 to the second quarter 1997. Share prices and economic indicators served as the variables (all measured as logarithmic four-term differences). We were unable to confirm a significant leading relationship for general price levels except for the CPI.[3] In testing for relationships with the

Results of Granger tests between share price and other variables

	Share prices --> Other variables	Other variables --> Share prices
CPI	* F value = 4.463	+ F value = 1.001
WPI	+ F value = 2.152	+ F value = 0.522
GDP deflator	+ F value = 2.442	+ F value = 0.461
Real GDP	** F value = 2.678	+ F value = 0.912
Real domestic private demand	* F value = 4.218	+ F value = 0.671
Real private-sector consumptive expenditures	** F value = 3.119	+ F value = 1.317
Real private-sector capital investment	* F value = 3.786	+ F value = 0.176

Note: * indicates significance at the level of 1%; ** indicates significance to the level of 5% and + no significance.

2 The Granger test was performed using a four-term lag model. The reason for selecting four terms (or, one year) was that our purpose was to verify the usefulness of share prices as an information variable for policy administration. Too long of a lead, even if it could be detected, would be of limited practical use. Obviously, however, it would be possible to arrive at analytical findings that differ from ours were the lag period changed.

3 Since foreign exchange rates and oil prices have an enormous impact on Japanese prices, we also performed a three-variable VAR Granger Test in which import prices, which directly reflect these movements, served as an exogenous variable. The results were not, however, significantly different.

real economy, we confirmed that share prices lead both real GDP and its component items (domestic private demand, private-sector consumptive expenditures, private-sector capital investment, etc.).

Time correlations

We next examined time correlations between share prices and other variables. We obtained the highest coefficient of correlation for real GDP and other real economic indicators for the full sample period, at approximately 0.5 with a lead of one year or less. We also divided the sample into smaller sub-periods (first quarter 1970 to fourth quarter 1974, first quarter 1975 to fourth quarter 1984, and first quarter 1985 to second quarter 1997). While the correlation was, for the most part, lost for the second sample period, the sub-sample with the smallest rate of share price change as shown by standard deviation, there is a clear correlation for the first and third sub-samples, both periods in which the rate of share price change was large. We would note, however, that the lead period for share prices differs considerably between the two sub-samples. Share prices, in other words, do lead the real economy, but the extent of the lead is uncertain.

Coefficients of time correlation between share prices and other variables

	1970Q1-97Q2	1970Q1-74Q4	1975Q1-84Q4	1985Q1-97Q2
CPI	0.334 (t = -7)	0.920 (t = -7)	-0.040 (t = -7)	0.133 (t = -8)
WPI	0.398 (t = -6)	0.881 (t = -6)	0.143 (t = -6)	0.251 (t = -8)
GDP deflator	0.358 (t = -7)	0.887 (t = -6)	-0.074 (t = -8)	0.219 (t = -8)
Real GDP	0.489 (t = -2)	0.777 (t = -1)	0.198 (t = -2)	0.737 (t = -8)
Real domestic demand	0.619 (t = -3)	0.794 (t = -2)	0.244 (t = -3)	0.705 (t = -8)
Real private-sector consumptive expenditures	0.462 (t = -1)	0.782 (t = -1)	0.117 (t = -4)	0.647 (t = -8)
Real private-sector capital investment	0.554 (t = -4)	0.954 (t = -3)	0.502 (t = -2)	0.781 (t = -8)
Standard deviation of rates of changes of share prices	22.02	32.17	9.59	23.80

Note: The coefficient of correlation is the largest from the t = -8 period to the t = +8 period (t < 0 indicates that share prices lead).

Using the Probit method to develop economic forecasts from share prices

We next used an econometric technique called the "Probit Method" to see if share prices were able to forecast two values of economic orientation (expansion or recession) even assuming that there is little set quantitative relationship between share prices and serial economic variables like GDP. The Probit method regresses the existence of an event (in this case, economic recession) back to a variable that is thought to have some relationship to the event (in this case, share prices), seeking the probability of an event's occurrence.[4] The results indicate some degree of usefulness (Figure 2) as share prices accurately predicted the economic recession of the first half of the nineties.

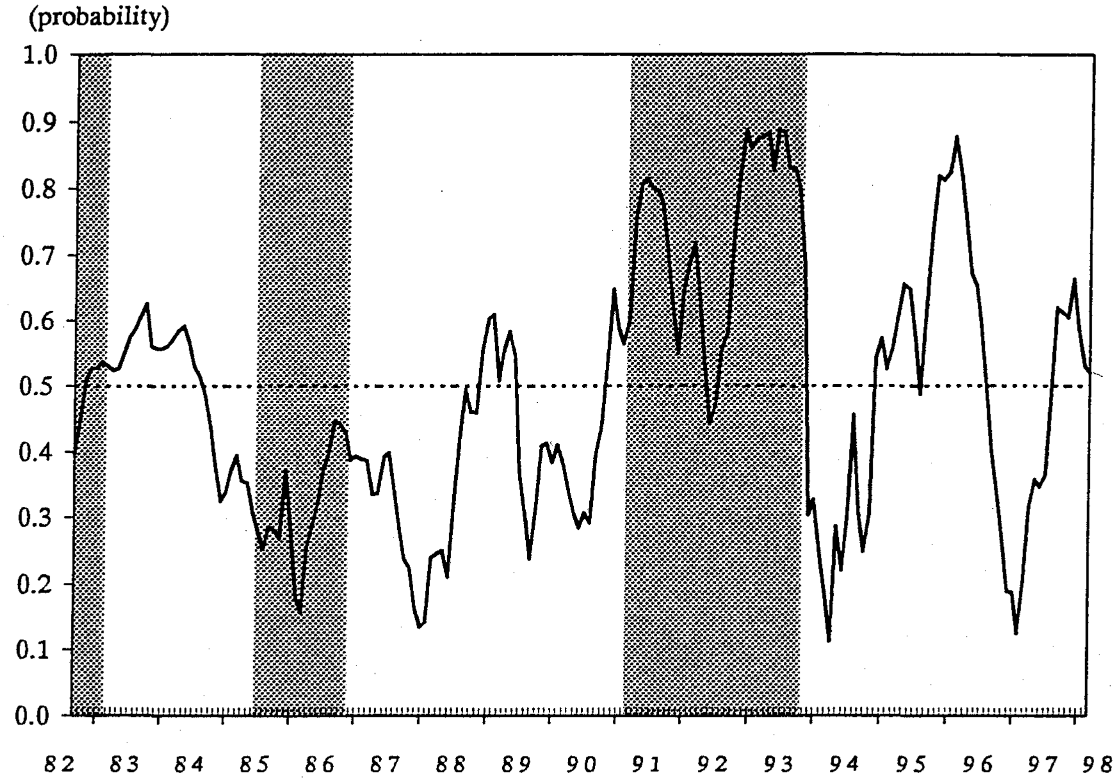
---

4 It is possible to consider the stock market as containing two kinds of participants, those who are "optimistic" about the economic future and those who are "pessimistic". Share prices reflect which group is stronger.

Figure 2

Predictive power of share prices using the Probit method

TOPIX (year-to-year, forecast 7 months later)



Method of calculation: First, we regress the variables (in this case, TOPIX) from period (t) to period ($t + x$), and forecast period ($t + x + k$) based on this regression. Next, we regress again from period (t) to period ($t + x$), and forecast period ($t + x + k + 1$). We repeat the procedure by shifting the estimation period one term ahead at a time. The purpose of this test is to confirm whether we can predict future recessions (out of sample period) by using the existing data (in the sample period).

Notes: The shaded areas show recessions based on the standard date of business cycles published by *Economic Planning Agency*. Each value shows the probability of recession calculated from data up to a specific number of months (in this case, seven) before the prediction period.

Source: Tokyo Stock Exchange.

2. Distinguishing characteristics of the formation of Japanese share prices

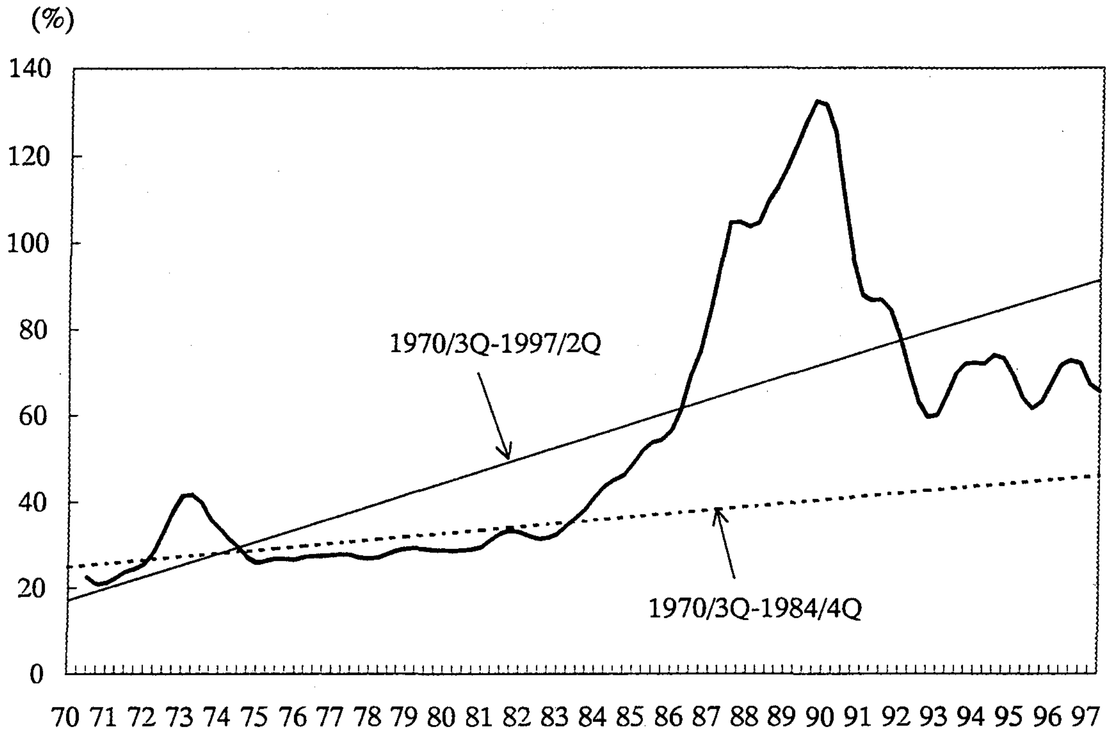
We have so far verified the usefulness of share prices as information variables for policy makers. This section focuses on the relationship between share prices and land prices as one of the distinguishing characteristics of past Japanese share price formation.

Relationship between market capitalisation and nominal GDP

We begin by looking at the long-term relationship between market capitalisation and nominal GDP (Figure 3). During the late eighties, the ratio of market capitalisation to GDP rose well beyond previous trend lines, but in the nineties it fell rapidly. This indicates the possibility that a bubble, which cannot be explained by any change in fundamentals, boomed and busted at this time.

Figure 3

The ratio of market capitalisation to nominal GDP



Notes: Gross market capitalisation consists of firms listed on the First Sections of the Tokyo Stock Exchange. Based on three-quarter moving average of end-month data. The lines indicate trends for each sample periods.

Sources: Tokyo Stock Exchange and Economic Planning Agency.

Relationship between the rate of share price change and the rate of land price change

The late eighties saw substantial increases in land prices, which indicate that the bubble formed across asset prices as a whole. When the relationship between the rate of share price change and the rate of land price change is considered over the short term, the two appear to move differently, in part because of the large swings in the rate of share price change (Figure 4, top). Over the medium to long term, however, their movements are similar. Indeed, the rate of land price change is virtually a backward moving average[5] of the rate of share price change (Figure 4, bottom). Theoretically, land prices and share prices should be formed by common macroeconomic factors like nominal GDP and interest rates, so it is rational that they would be linked. However, it appears that the correlations between share prices and land prices are particularly strong in the case of Japan.

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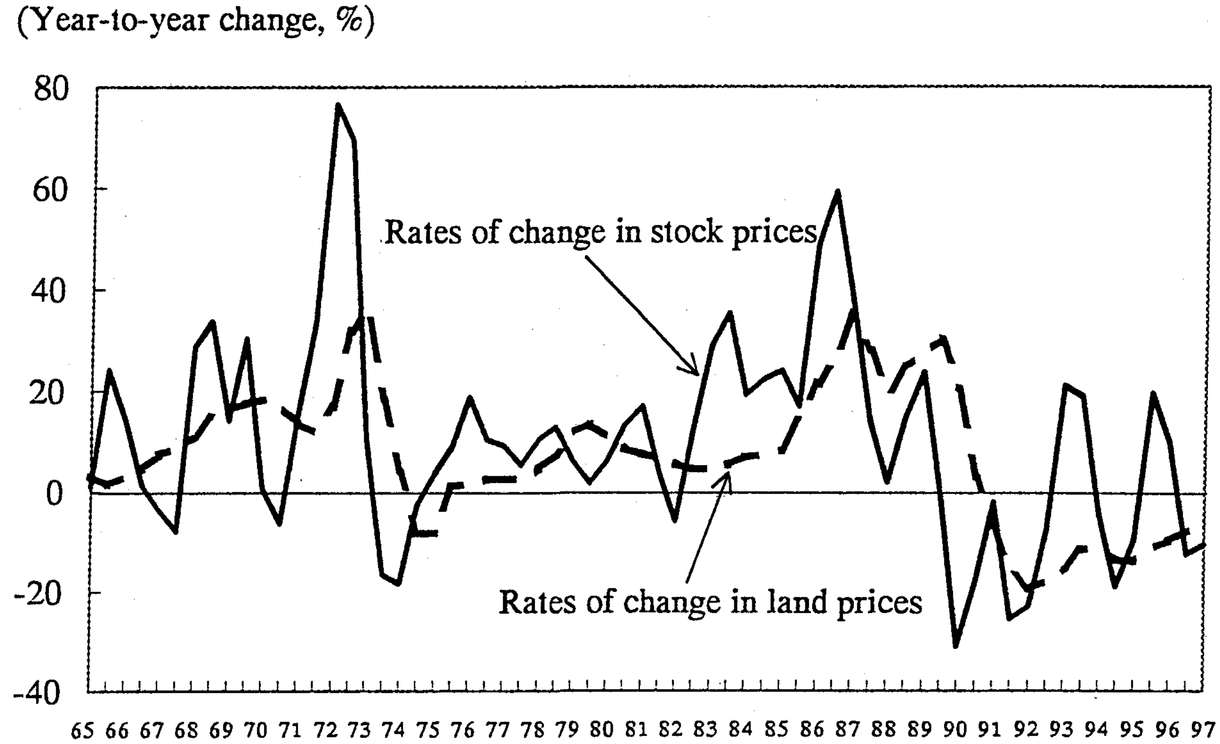
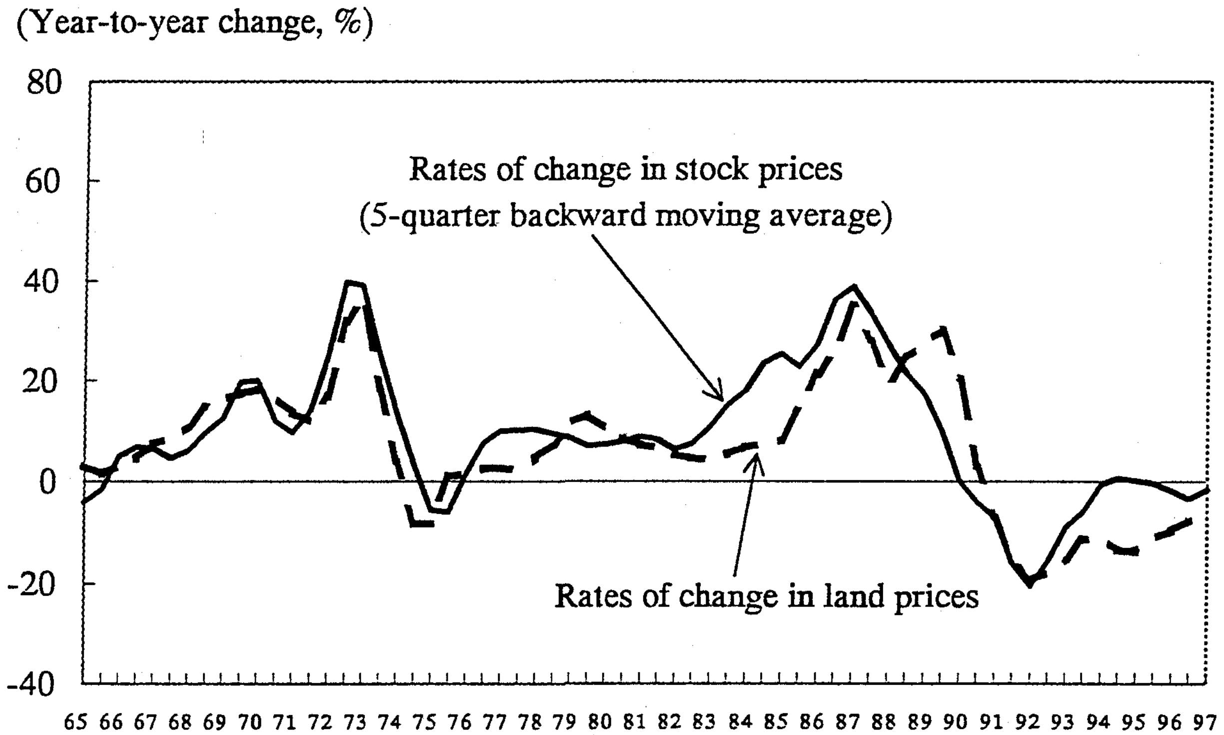
[5] The reason for a "backward" rather than a "median" moving average is probably that land lacks liquidity and the land market therefore tends to react more slowly to changes in the environment than the stock market. From a technical standpoint, we would also note that there is an even longer lag required before prevailing market prices are reflected in land price indexes.

Figure 4

Changes in land and stock prices





Notes: Urban land price index (six major cities, average of all uses) used for land prices. It is assumed that the trend change in land prices in 1996H2 would continue in 1997. TOPIX used for stock prices. Figures for both land and stock prices are six-month data for April-September and October-March.

Sources: Tokyo Stock Exchange and Japan Real Estate Institute.

Relationship between corporate net asset values and market capitalisation

Japanese accounting standards do not use market values to appraise assets, so it is difficult to measure corporate net asset values in terms of the reacquisition cost (the market price), but a macroeconomic approximation can be made if a few assumptions are allowed (Figure 5, top). In the late eighties, rising asset prices drove up the value of land and shares owned by companies, which in turn caused a rapid increase in corporate net asset values. When this trend is overlaid on the trend lines for market capitalisation, an almost exact match is discovered (Figure 5, bottom). This indicates that the stock market of the late eighties valued the rise in corporate land and share assets (including unrealised gains) virtually without modification. As long as the market price of corporate assets provides an accurate reflection of the profitability of the asset – in other words, as long as it is close to the discounted present value of the profits that the asset will produce in the future – then it is natural that a change in the market price of an asset will be reflected in the market capitalisation of a company holding the asset. It is possible, however, that the stock market of the late eighties was valuing assets with the bubble that had formed in land prices.[6] We can assume that a mechanism then took root in which share prices valued in terms of rising land prices further boosted the value of the shares issued by companies that had extensive stock portfolios because of crossholding relationships.[7] If that was indeed the case, when the bubble burst and land prices began a sustained decline in the nineties, the reverse mechanism took root.[8]

The reasons behind strong ties between share prices and land prices

The discussion above should make it clear that the ties between share prices and land prices in Japan are far stronger than what would be expected from a general price arbitrage relationship between different classes of assets. That begs the question of why such linkage would exist, a question that is difficult to answer quantitatively, but which can be qualitatively addressed by the following points.

First, during the postwar reconstruction and high growth period, the price of both shares and land kept rising and both assets were used as a means of diversifying investments. As a result, there is a very strong arbitrage relationship between their prices. Although its profitability varied significantly, land has, in general, been considered an advantageous asset to hold, in part because of

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6 In addition, an increase in the unrealised gains that is unlikely to lead to an increase in future cash flows – say, unrealised gains on land that the company is using for production activities or idle land that the company has no plans to use – should not be reflected in the share price at all, except if the company is an M&A target (in which case, the unrealised gains would be realised in the form of cash flow). During the late eighties, there were many attempts to justify share-price levels using the "Q ratio" (Market capitalisation/Gross market valuation of the company's assets – Gross liabilities) or market priced PBR. With hindsight, these can only be termed misleading. Such justifications confuse the theoretical breakup value of the company with its value as a going concern that produces revenues in the form of cash flow. By rights, the only assets that should be counted for such valuations is capital equipment. Similarly, Tobin's Q is an index of corporate strength in relation to asset holdings that takes share prices as a given, not an evaluation of share prices themselves (in other words, the idea has been reversed). Moreover, these theories and indexes have even less usefulness in cases like those currently being debated in the United States in which software and other intangible assets are not accurately measured in corporate accounts.

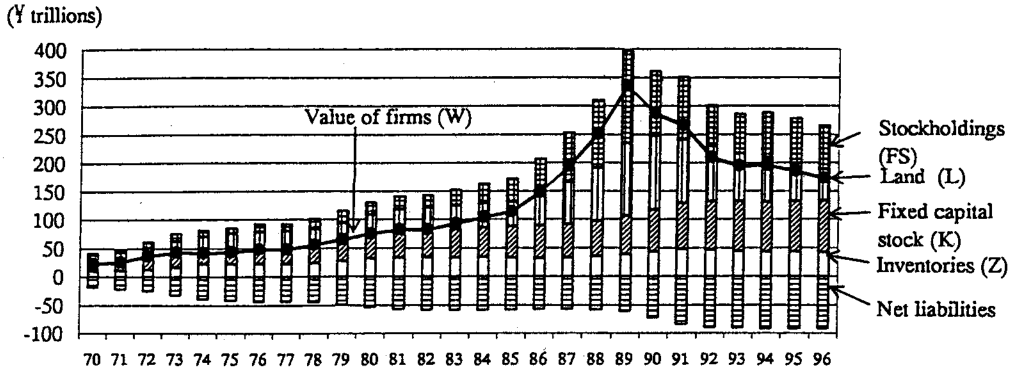
7 Taking the crossholding ratio as α, then a rise in the value of corporate assets other than shares (= 1) would have the effect of increasing the market capitalisation of the sector as a whole by $1 + \alpha + \alpha^2 + \alpha^3 + \ldots\ldots = 1 / (1 - \alpha)$.

8 However, the timing of the market's downturn indicates that share prices were the leader. Share prices turned in 1990 and land prices not until 1991. What probably happened was that the highly liquid stock market was quicker to react to the increased risk of land price drops brought by changes in macroeconomic conditions (higher long-term interest rates), and government moves to clamp down on land prices (the imposition of regulations on total lending to the real estate industry).

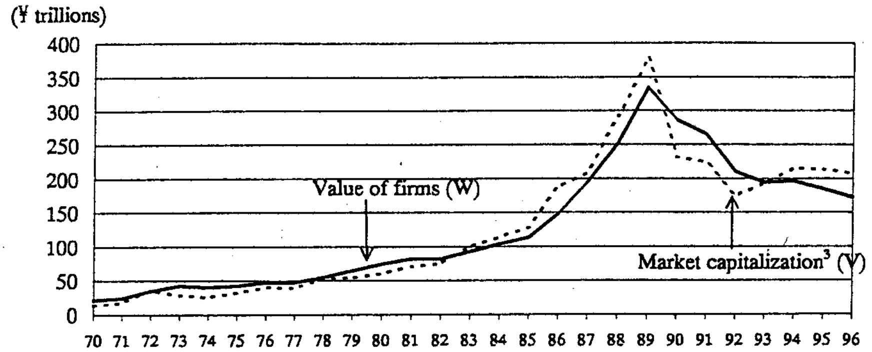
Figure 5

Assessment of the value of firms[1]

Value of firms in terms of replacement cost[2]



Value of firms in terms of replacement cost and market capitalisation



[1] Figures are estimated by the Bank of Japan, based on firms listed on the First and Second Sections of the Tokyo Stock Exchange excluding banking, insurance, securities and other financial services industries.

[2] Value of firms in terms of replacement cost (W) = *net asset value of firms* = $K + Z + L + FS + FA - B$, where:

K = fixed capital stock - land value (at book value) (aggregate number of firms listed on the First and Second Sections of the Tokyo Stock Exchange according to NEEDS),

Z = gross value of inventories (according to NEEDS),

L = value of land at market value. For 1995 and earlier figures, derived by multiplying the non-reproducible tangible asset/cash and deposits ratio, in the "non-financial corporate enterprises" sector in Annual Reports on National Accounts, by cash and deposits according to NEEDS. For 1996 figures, calculations based on year-to-year changes in the urban land price index (six major cities, average of all uses),

FS = total value of stockholdings at market value = stockholding at book value (securities holding (according to NEEDS) x ratio of stock/securities (according to Financial Statements of Incorporated Business, Quarterly)) + unrealised gains on securities held by firms (according to Shuyo Kigyo Keiei Bunseki and NEEDS) for 1994 figures. For 1993 figures and earlier, calculations based on year-to-year changes in stockholdings at market value in the "non-financial corporate enterprises" sector in Annual Reports on National Accounts. 1995 and 1996 figures based on year-to-year changes in Market Capitalisation,

FA = financial assets (excluding stockholdings) = gross value of assets (according to NEEDS) - fixed capital stock (K) - inventories (Z) - land (at book value) - stockholdings (at book value),

B = gross liabilities (according to NEEDS)
Net liabilities - gross liabilities (B) - financial assets (excluding stockholdings) (FA).

[3] Market capitalisation = (gross value of stock of firms listed on the First and Second Sections of the Tokyo Stock Exchange at market value - gross value of stocks of banking, insurance, securities and other financial services industries at market value) / number of listed firms x number of sample firms in NEEDS.

Sources: Economic Planning Agency, *Annual Reports on National Accounts*, Ministry of Finance, *Financial Statements of Incorporated Business, Quarterly*, Bank of Japan, *Shuyo Kigyo Keiei Bunseki* (*Analysis of Financial Statements of Principal Enterprises*), and Nihon Keizai Shimbun, Inc., *NEEDS* (*Nikkei Economic Electronic Databank System*).

regulatory factors (the tax code and land use regulations).[9] The result has been to obfuscate the price formation standards for land and has kept land prices rising at the same rate as share prices.

Second, in the postwar period lending has generally been secured with real estate. This has induced a process where rising land prices increase corporate fund-raising abilities, which in turn spurs an expansion in capital investment and corporate profits and subsequently translates into higher share prices.

Third, share crossholding arrangements between companies have reinforced the linkage between land and share prices by encouraging the stock market to value companies in terms of their net assets.

In Section 1 we confirmed the usefulness of share prices as a predictor of real economic activities. As far as the Granger test results and time series correlations show, share prices lead particularly strongly such component items in real GDP as domestic private sector demand and private sector capital investment. Also, in another paper using Granger tests and time series correlation analyses to verify the leading relationship of land prices to real economic indicators, we obtained the same results as for the share prices.[10] Therefore, share prices and land prices have a strong relationship and are probably both useful as an information variable for real economic activities.[11]

3. Share price valuation using the framework of a dividend discount model

In Section 2 we worked from the assumption that the stock market assesses corporate net asset values and went on to consider the formation of share prices since the bubble. In this section, we analyse share price formation using the framework of a "dividend discount model", which expresses share prices as the present value of the dividends (or the profits that are their source) produced by the company in the future. More specifically, we will use the fact that the yield spread (long-term interest rates – earnings yield), which is often employed as a standard for valuing share prices in relation to interest rates, is equal to the difference between the expected growth rate for nominal corporate earnings minus the risk premium to examine the factors behind the recent share price slump in terms of these two measures. Below is an outline of the framework used.

We will assume that current nominal earnings per share (E) increase year to year by a fixed growth rate (g). We can therefore use the following formula to calculate the present value (P, equal to the share price) of the stream of future earnings discounted for the rate of yield demanded by investors (δ).

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9 The effective rates of both the inheritance and the land-holding taxes were kept extremely low. In addition, land-use regulations were often administered ambiguously, which allowed, for example, prices to form for agricultural land on the assumption that it could be converted to residential or commercial use. We must also note the influence of the postwar "land myth" (that "you will never lose by owning land" or that "land is the most advantageous asset to invest in").

10 The basis for share prices, corporate profitability, is strongly influenced by foreign demand, whereas the basis for land prices, rent, is assumed to depend on private sector domestic demand. Therefore, it is natural that land prices are useful as an information variable, leading, particularly, private sector domestic demand.

11 Asset prices lead the real economy not only because market expectations anticipate future changes in macroeconomic conditions, but conceivably also because changes in share prices themselves exert a direct influence on demand and spending through the wealth effect on households and changes in the cost and availability of corporate funding. However, our purpose in this paper is not to discuss the transmission mechanism between share prices, land prices, and real economic activities. Hence, we refrain from delving any further into these issues here.

$$P = \frac{E}{1+\delta} + \frac{E(1+g)}{(1+\delta)^2} + \frac{E(1+g)^2}{(1+\delta)^3} + \cdots\cdots$$

$$= \sum_{n=1}^{\infty} E(1+g)^{n-1}/(1+\delta)^n = \frac{E}{\delta - g}$$

$$\therefore \frac{E}{P} (= \textit{Earnings per share / price} < \textit{Earnings yield} >) = \delta - g$$

The rate of yield demanded by investors (δ) will probably be the long-term interest rate (r) plus a risk premium (ρ), so that:

$$\frac{E}{P} = \delta - g = r + \rho - g$$

$$\therefore \textit{Yield spread} = r - \frac{E}{P} = g - \rho$$

The price/earnings ratio and the yield spread

Price/earnings ratios since the late eighties show large upwards and downwards shifts, with peaks coming in 1987 and 1994. Current levels are about average (Figure 6, top). Evaluations of share price and price/earnings ratio levels must, however, take account of the correlations with interest rates levels. The yield spread is the difference between long-term interest rates and the earnings yield, which is the inverse of the price/earnings ratio. Trends show that an average line of 3.5% held until the early nineties, but since 1995 the yield spread has moved substantially downwards and share prices would, superficially, appear to be "cheap" (Figure 6, bottom). As we have already discussed, a contraction in the yield spread would, in theory, indicate a decline in expected earnings growth rates or an expansion in the risk premium, or perhaps both. These factors must be taken into account when evaluating current share price levels. In the pages that follow, we consider the background to changes in the yield spread in some detail, but before doing that we must make two adjustments to the price/earnings ratio in order to more accurately capture yield spread levels.

The first adjustment is to correct for the influence of share crossholding arrangements (see Appendix A for the correction method). Share crossholdings have no direct impact on corporate profitability and so, in theory, do not affect share prices.[12] However, they are generally thought to have the effect of raising the apparent price/earnings ratio.

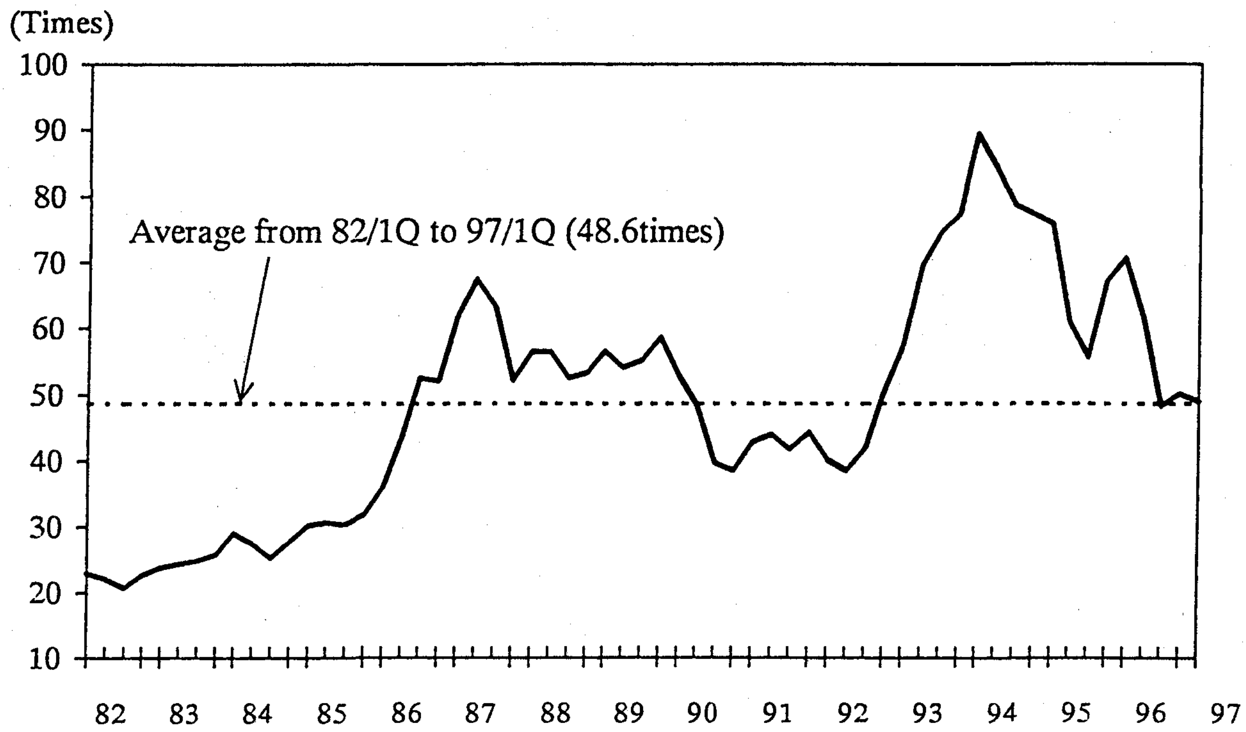
The second adjustment is to correct for business cycles (see Appendix B for the correction method). If we assume that near-term corporate profits will undergo large swings because of the business cycle, but that the expected growth rate for nominal earnings remains constant, then when the market predicts the stream of future earnings, the present value of earnings will differ from actual earnings and will be closer to the trend line. Therefore, if the economy is currently in recession and the markets expect corporate earnings to recover in the future, the price/earnings ratio will be on the high side. Likewise, if the economy is currently robust but the markets expect corporate earnings to decline in the future, then the price/earnings ratio will be on the low side.

---

12 Share crossholdings between companies have no impact on the actual value of a company because the increase in dividend income that comes from the shares that a company holds will be offset by dividends paid out to companies that hold its shares. This can be verified from a simple numerical example. However, crossholdings and their unwinding may have a short-term impact on share prices via the supply and demand mechanism, and this will be more the case the greater the incompleteness of the market and the asymmetry of information among participants.

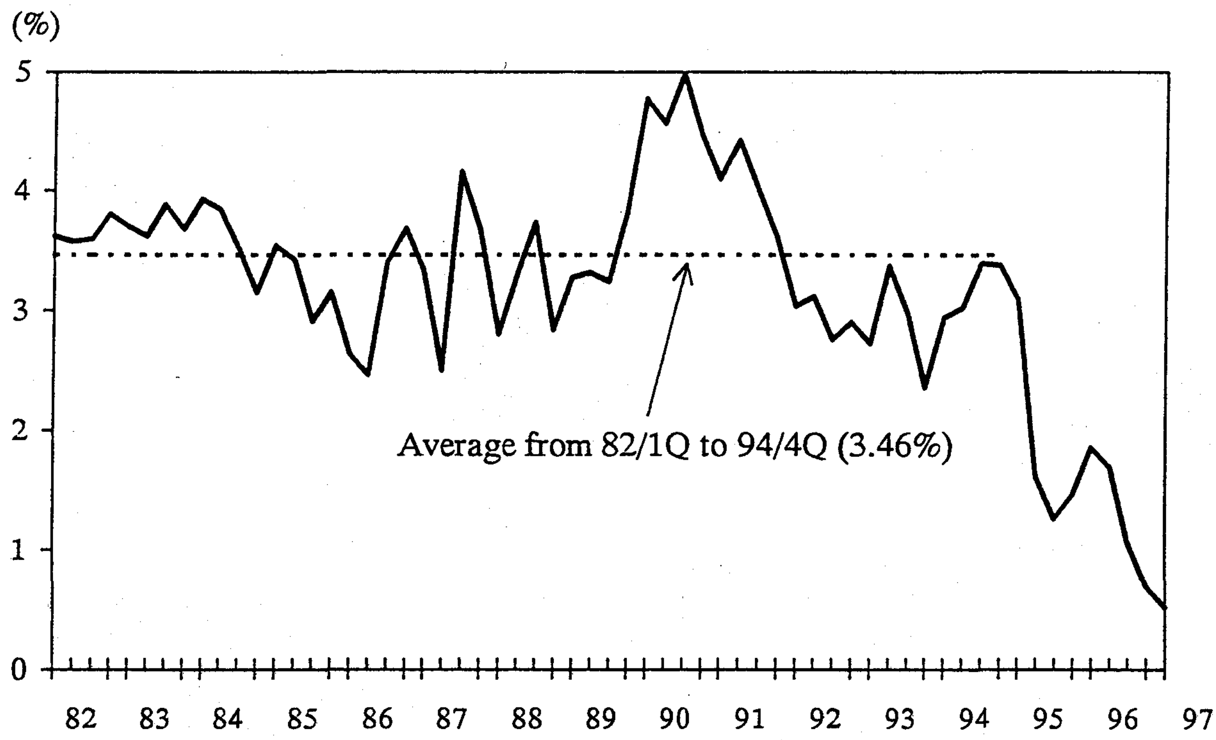
Figure 6

Price/earnings ratio



Source: Daiwa Research Institute.

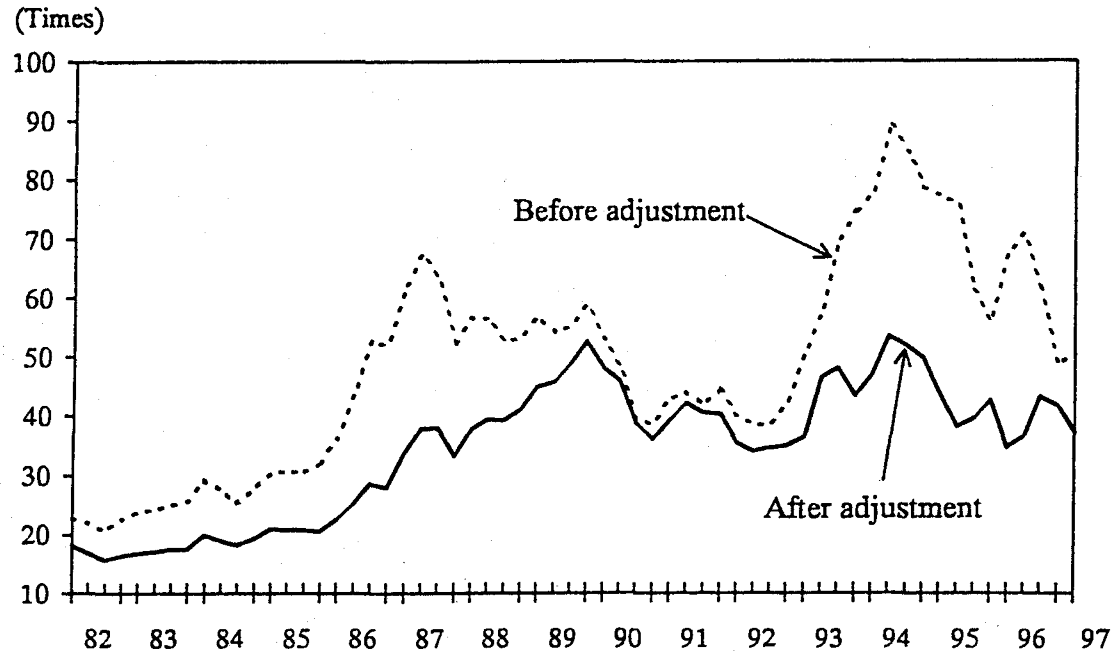
Yield spread



Notes: Yield spread = yield on government bonds (10-year) - earnings yield = Expected growth rate of firms' nominal earnings - risk premium in stock markets. Data for "Banks" are excluded from 1996Q1 and Q2.

Figure 7

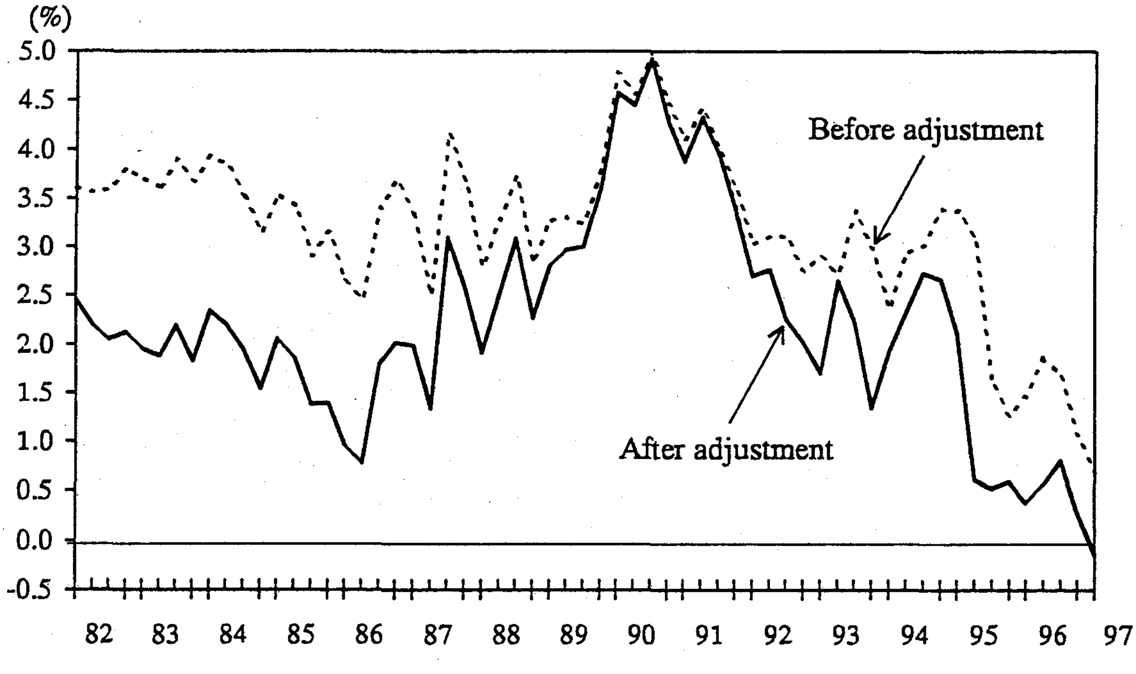
Adjusted price/earnings ratio



Method of calculations: *PER* after adjusting for the firms' share crossholding factor = *PER* x (1 - θ) / (1 - *D* x θ), where θ is the share crossholding ratio and *D* the payout ratio. *PER* after adjusting for the firms' share crossholding factor and business cycles factor = Coefficient for adjusting business cycles factor x *PER* after adjusting for the firms' share crossholding factor.

Notes: The data for "Banks" are excluded from 1996Q1 and Q2. *PER* is expected *PER* based on the survey by Daiwa Research Institute. The payout ratio is from the National Conference of Stock Exchanges and the share crossholding ratio is from Daiwa Research Institute. In adjusting for the business cycles factor, we used the GDP gap.

Adjusted yield spread



Notes: Yield spread = yield on government bonds (10-year) - earnings yield = Expected growth rate of firms' nominal earnings - risk premium in stock markets. Data for "Banks" are excluded from 1996Q1 and Q2.

When the price/earnings ratio is adjusted for both these factors, which levels appear lower than unadjusted price/earnings ratio[13] (Figure 7, top). Similarly, the fluctuations seen since the late eighties are smoothed out.

We are now ready to use the corrected price/earnings ratio to trace the yield spread (Figure 7, bottom). One can see that it rose rapidly in the late eighties and declined rapidly after 1991. This paints a much clearer picture of the changes in share price levels in relation to interest rate levels during the formation and collapse of the bubble.[14]

Expected growth rate of nominal earnings and the risk premium

We will calculate the risk premium using the yield spread and assuming a constant expected growth rate for nominal earnings below. Being a remainder, the risk premium will obviously change somewhat according to assumptions about the expected growth rate of nominal earnings, so results must be viewed with a certain degree of latitude. Even so, attempts such as ours are useful in viewing share price formation trends over the medium term.

For the expected growth rate of nominal earnings we use the medium-term real corporate growth rate forecasts found in the *Survey of Corporate Activities* published by the Economic Planning Agency.[15] From this base we add an expected CPI inflation rate[16] derived from an adaptive expectations model, thus obtaining a closer approximation. The expected growth rate of nominal earnings thus obtained was over 9% at the end of 1982, but during the eighties, it declined to the 3% level before turning upwards again in the early nineties. At the end of 1991, it stood in the 6% range. It has again undergone a decline and is currently in the 2% range (Figure 8, top).

The next step is to use the adjusted yield spread and the figures for the expected growth rate of nominal earnings to derive the risk premium observed in the markets. Our findings indicate that the risk premium declined rapidly in the late eighties and was at one point close to zero before rising rapidly in the nineties, peaking in 1992, declining through 1994, and then turning upwards again in 1995 (Figure 8, bottom). In as much as it is calculated after the fact based on the expected growth rate of nominal earnings and several other assumptions, this risk premium should be viewed as a "balance" in which are subsumed the swings in market expectations and mistakes in market forecasts. It is hard to consider it an accurate measure of the risk premium included in the a priori rate of return demanded by investors (this is the same as the discount rate in the dividend discount model). In fact, it is likely that the rapid decline in the risk premium at the end of the eighties reflected the stock price bubble (a stock price movement that departs from fundamentals).

---

[13] Even the corrected levels show price/earnings multiples of about 35, which are high in comparison to the United States (the S&P 500 has a multiple of about 20). When the price/earnings ratio is used to make international comparisons between markets, differences in statutory reserve requirements, fixed asset depreciation, and other corporate accounting practices must be taken into account above and beyond interest rate levels. One cannot simply conclude on the basis of the price/earnings ratio that a market is "dear" or "cheap". Other things being equal, the price/earnings ratio will be higher the lower interest rate levels go. Some analyses also indicate that the price/earnings ratio of Japanese companies would be considerably lower were US-style accounting practices used.

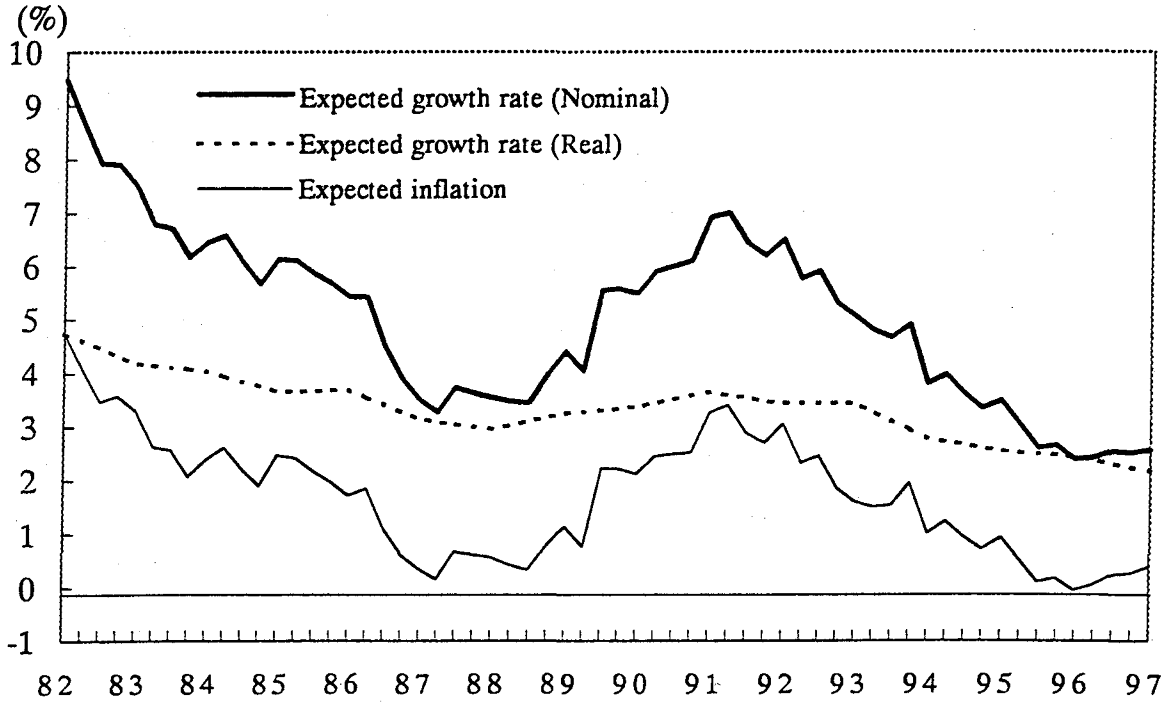
[14] A bubble is a price movement that departs from fundamentals. When a price movement containing a bubble is later explained in terms of a fundamentals model, the yield spread and the risk premium (discussed later) are likely to show "excessive" swings.

[15] Ideally, the expected growth rate would be a measure of investor expectations, which, if one assumes information to be asymmetric, may not match the expected growth rate of the companies themselves. However, data constraints force us to use the values from the corporate survey.

[16] Approximated with a lagged eight-term moving average of term-to-term CPI growth.

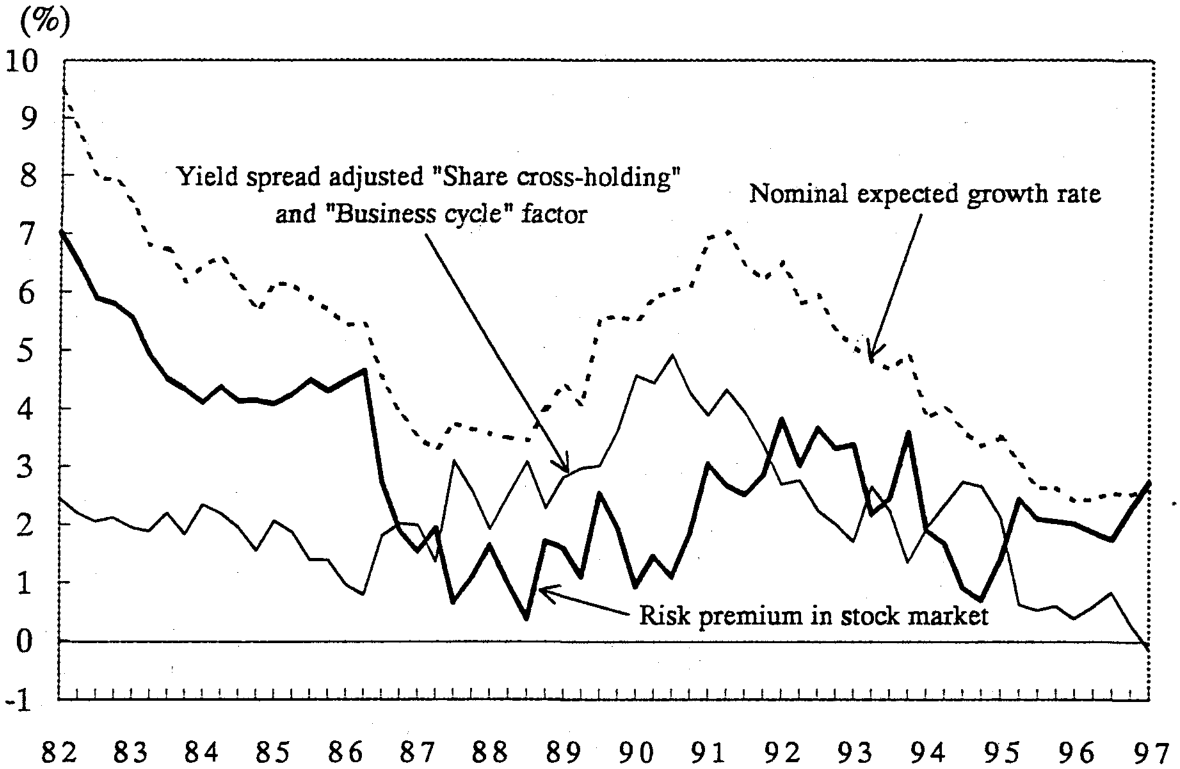
Figure 8

Expected growth rate of firms' nominal earnings



Notes: Expected growth of firms' real earnings is based on Economic Planning Agency, *Corporate Behavior Survey*. Expected inflation is calculated from an Adaptive Expectations Model.

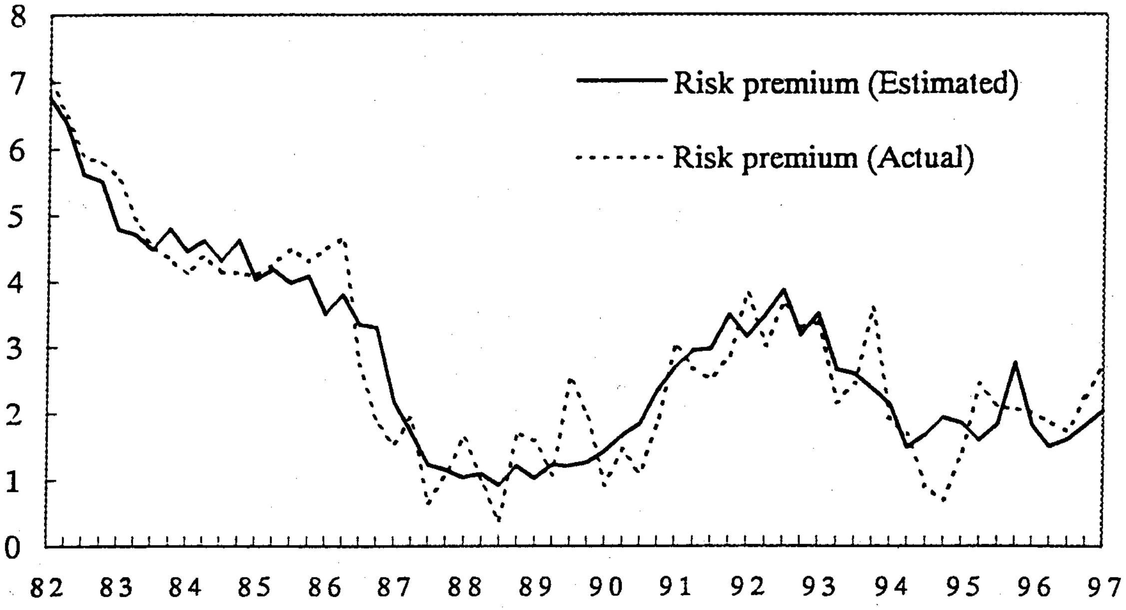
Risk premium in the stock market



Notes: *RP* = Risk premium in stock market = *g* - *YS*, where *g* = Expected growth rate of firms' nominal earnings and *YS* = Yield spread adjusted "Share crossholding" and "Business cycle" factor.

Figure 9

Theoretical risk premium in the stock market

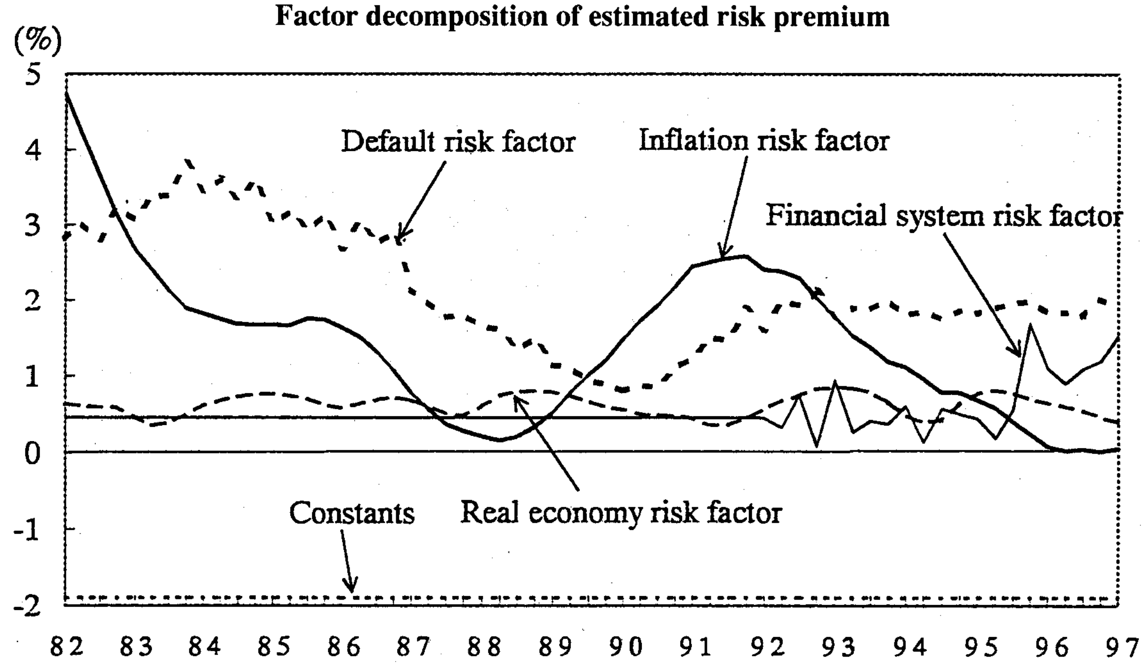


Method of calculation: Risk premium in stock market = -1.89 + 0.81 x inflation risk factor +0.18 x real economy risk factor
(-3.9) (11.4) (1.7)
+ 4.22 x Default risk factor + 6.03 x Financial system risk factor
(9.8) (3.4) ( ) = t-value

Sample period = 1984Q1-1997Q1
Adjusted R-square = 0.860 S.E. = 0.586 D.W. = 1.512

Variables: Inflation risk factor: 8-quarter backward moving average of year-to-year change in CPI; real economy risk factor: 12-quarter backward standard deviations of year-to-year change of IIP; default risk factor: bankruptcy rate of firms; financial system risk factor: CD(3M) - TB(3M).

Factor decomposition of estimated risk premium



In other words, there are many problems in the estimation of the risk premium. Nonetheless, we have regressed the risk premium that we calculated on other variables considered likely to influence the risk premium. This was done because it provides a means for exploring the factors behind share price formation using the framework of the fundamentals model (Figure 9, top). As proxy variables for earnings and interest rate fluctuation risks, we used the CPI and industrial production; as proxy variables for credit risk, we used the corporate bankruptcy rate and the CD-TB rate spread, giving us a total of four variables.[17] Our estimates indicate that the decline in the inflation (CPI) and default risks (corporate bankruptcy rate) contributed to the decline in the risk premium seen in the late eighties (Figure 9, bottom). In the nineties, both of these variables rose, which caused the risk premium to rise. Then in the mid-nineties, inflation risk again declined, but default risk remained high and financial system risk (the CD-TB rate) rose, limiting the declines in the risk premium.

The current risk premium is in fact at lower levels than it was in the early eighties, which could indicate that there is still room for the risk premium to rise (and therefore for share prices to decline). Certainly, the bankruptcies of medium-sized constructors illustrate that as long as the "negative" legacy from the bubble continues, there will be room for the premium against credit risk to expand. Nonetheless, the expected inflation rate has vastly declined from what it was in the early eighties, and if the markets interpret this as meaning that there is little risk of a large rise in long-term interest rates, it would not necessarily be irrational for the risk premium as a whole to be lower than the levels of the early eighties.

Factors in the Japanese share price slump

Let us turn once again to the dividend discount model and re-examine the factors at work in share price formation.

1) In the early eighties both the price/earnings ratio and the yield spread were stable (when both are adjusted for share crossholding and cyclical factors, and so throughout). During this period, both the expected growth rate of nominal earnings and the risk premium declined.

2) In the late eighties, both the price/earnings ratio and the yield spread rose. During this period, the expected growth rate of nominal earnings rose, while the risk premium remained low.

3) In the nineties, the price/earnings ratio remained at roughly the average levels of the late eighties, but the yield spread consistently declined. During this period, the expected growth rate of nominal earnings declined and the risk premium rose.

The question is then how to view this analysis in light of the relationship between share prices and land prices – the high probability that during the late eighties, the stock market valuation of corporate net assets took at face value the rise in land prices, which itself contained a bubble.[18]

It can be said that the rapid decline in the risk premium during the late eighties corresponded with the bubble portion of land price valuation. Indeed, if the risk premium is explained in terms of a model that regresses all variables, then a bubble-inspired rise in land prices will be observed as a decline in credit risk. In the late eighties, the default risk (corporate bankruptcy rate)

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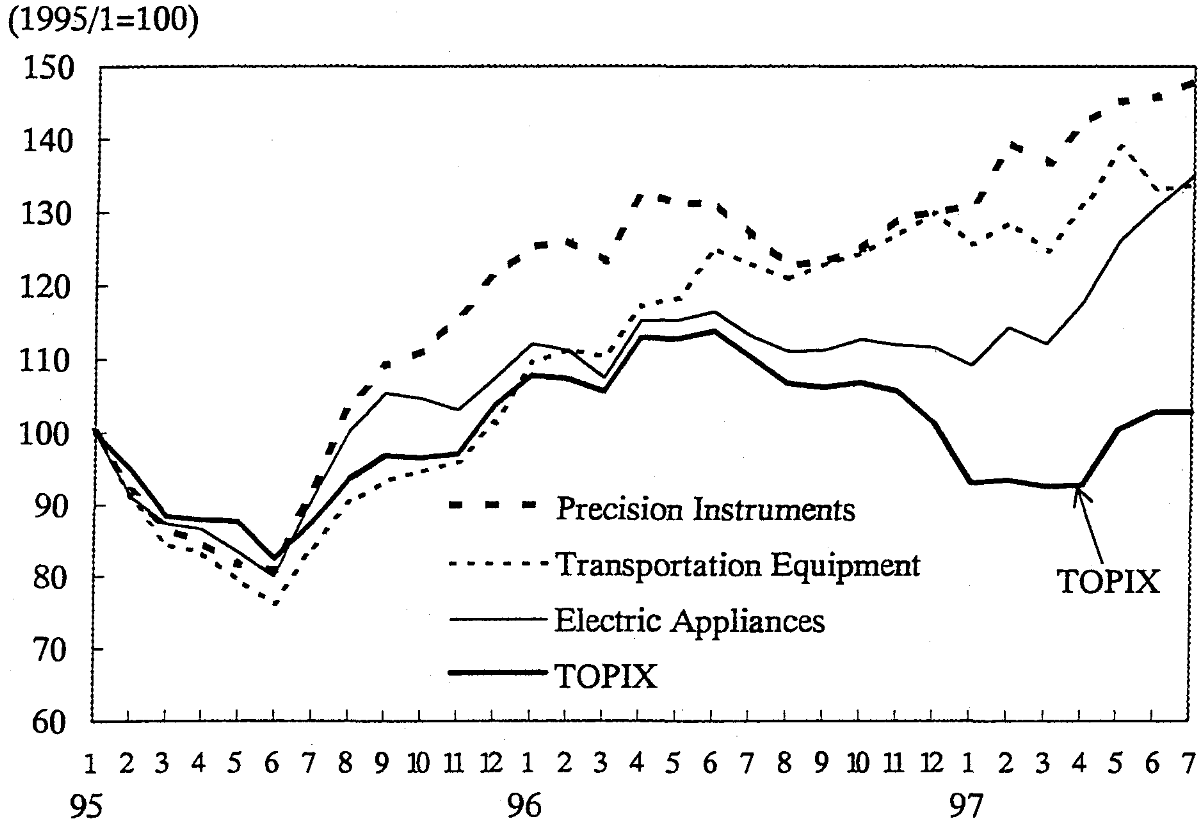
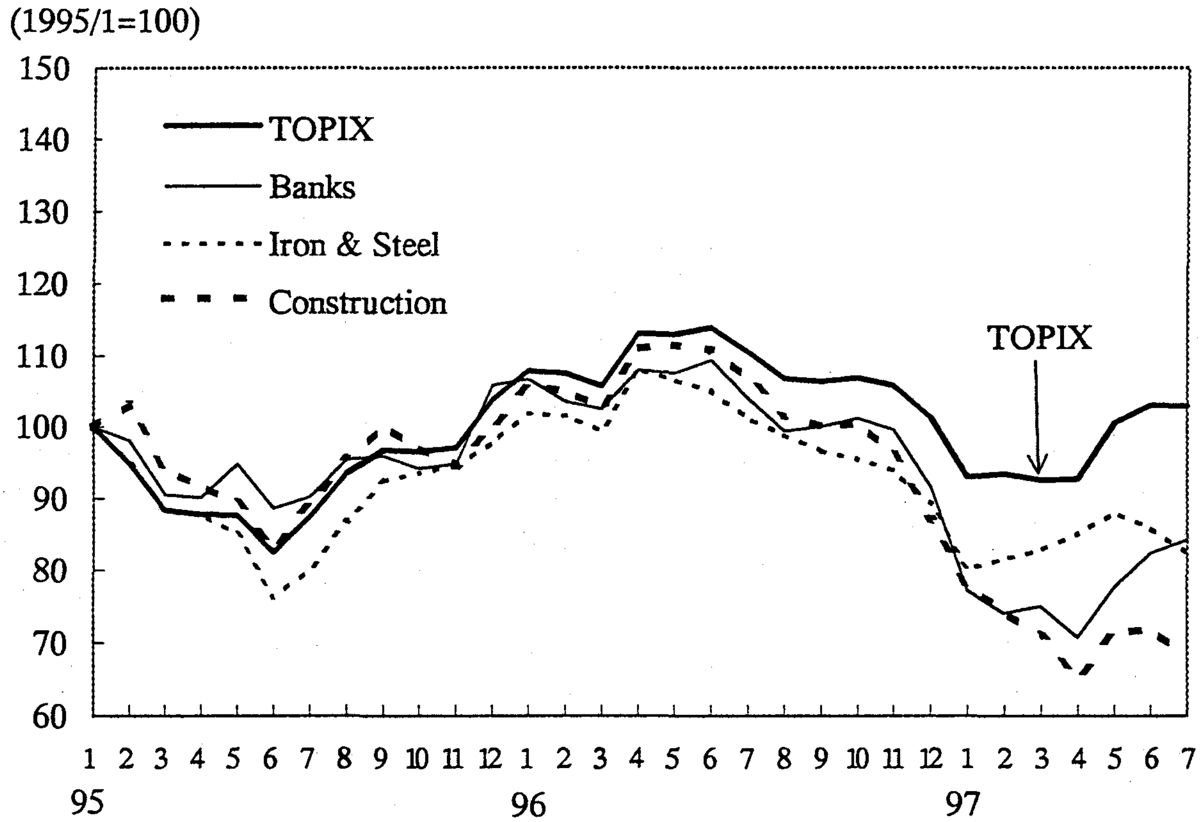
[17] We added share-price volatility as an explanatory variable to serve as a proxy for price-fluctuation risks, but this had no significance. Industrial production (standard deviation from the previous year) may be considered a proxy for price fluctuation risk in this regression.

[18] In the late eighties, it was evident that low interest rates and high expected growth of nominal earnings acted to push up both land and share prices. The issue here is how to comprehend, within the framework of Dividend Discount Model, the fact that the stock market was influenced by the land price bubble which is thought to be included in the increase in the value of firms' net assets.

Figure 10

Stock price index by industry





Source: Tokyo Stock Exchange.

declined sharply, and it is conceivable that a major part of this was the fact that rising land prices produced a rise in the collateral value of corporate assets. In other words, if one can assume that land prices will continue to rise or at least not decline, then it is probable that the stock market risk premium declined. Coming into the nineties, however, the reverse phenomenon was observed as land prices went into decline. Within the context of the dividend discount model, the decline in market capitalisation (the decline in share prices) that occurred almost in parallel with the decline in corporate net asset values caused by falling land prices can be captured as a rise in the risk premium due to higher credit risks. Since the mid-nineties, the default risk, as measured by the corporate bankruptcy rate, has been flat, but the financial system risk, as measured by the spread between the CD and TB rates, has risen, which has caused credit risk as a whole to rise. Thus the basic factor in the rapid decline in the yield spread during the nineties was the decline in the expected growth rate of nominal earnings, though the increase in the credit risk premium caused by falling land prices also played a role. This is what resulted in a slump in Japanese share prices in contrast to the booming markets in other industrialised countries.

We would note in conjunction with this that while share prices as a whole have been slumping in recent years, those for electric and precision equipment companies, which as far as corporate earnings and the risk premium go are less vulnerable to the impact of falling land prices, have been comparatively strong (Figure 10, top). Likewise, sectors like banking and construction that are very vulnerable to the effects of land price drops have seen major declines in their share prices (Figure 10, bottom). In other words, there has been contrasting developments among share prices.

4. Structural changes in the stock market

In the previous section we examined the factors behind the slump in Japanese share prices that has prevailed through most of the nineties, finding that it matched trends in macroeconomic factors, for example, the decline in the expected growth rate of nominal earnings and the drop in land prices. During this period several phenomena were observed in the stock market which seemed to augur changes in the market's structure. While it is not clear at this point what impact these phenomena have had on share prices, they do provide a wealth of hints about how to observe the stock market and share prices in the future, so they are described briefly in this section.

Changes in investors

Among the most pronounced changes in the stock market is the increased weight of foreign investors as players in the market. We divided investors into financial institutions, industrial corporations, personal investors, and foreigners, and charted their share of trading (by value) for the last ten years. In the late eighties, foreigners accounted for 11.5% of trading, but by the mid-nineties their share had soared to 27.8%, and in the first half of 1997 they have been responsible for 34.4% of trading, fully one-third of the money changing hands. On the other hand, the share of personal investors fell by half (15.9% in the first half of 1997) from 31.2% in the late eighties. The personal investors' separation from the stock market was probably caused by the after-effects of losses suffered when the bubble burst as well as the intensification of distrust in the stock market from repeated scandals of security companies.

Turning to the percentage of shares owned by different sectors, we find that the weight of personal investors declined between the end of 1985 (FY) and the end of 1990 (FY), while that of financial institutions and industrial corporations rose. Between the end of 1990 (FY) and the end of 1996 (FY), the weight of personal investors was flat, that of financial institutions and industrial corporations declined, and the weight of foreign investors rose sharply from 4.2 to 9.8%.

Share of trading (by value; %)

	1988-90	1991-93	1994-96	January-June 1997
Financial institutions	38.0	36.9	38.8	41.5
Industrial corporations	14.9	9.2	6.7	5.2
Personal investors	31.2	27.8	23.0	15.9
Foreigners	11.5	21.6	27.8	34.4
Others	4.4	4.5	3.7	3.0

Note: Totals for trading on the First and Second Sections of the Tokyo, Osaka, and Nagoya markets. Financial institutions include investment trusts, pensions, life insurance companies and other institutional investors.

Breakdown of share ownership

	End of FY 1985	End of FY 1990	End of FY 1996
Financial institutions	42.2	45.2	41.3
Industrial corporations	24.1	25.2	23.8
Personal investors	25.2	23.1	23.6
Foreigners	5.7	4.2	9.8
Others	2.8	2.3	1.5

Source: National Securities Exchange Council, *Report on the Survey of Share Distribution.*

Unwinding of share crossholding relationships

One of the reasons that ownership proportions have changed is probably the unwinding of share crossholding arrangements. During the late eighties, the percentage of shares in the portfolios of financial institutions and industrial corporations rose, in part because corporations and institutions took advantage of the rising share prices of this period to increase their capital, and some of the new shares issued were underwritten by other companies and institutions as part of crossholding arrangements. In the nineties, this has changed. Companies are unwinding their crossholding arrangements, and many of the shares involved are being picked up by foreigners, whose ownership percentage has increased by a corresponding amount. It is not just foreigners who have bought the shares being released; they are also going to pension funds, which are included among "financial institutions" in our statistics. At the end of 1990 (FY), pension funds owned only 0.9% of the stock in Japan, but by the end of 1996 (FY) their share had increased to 2.3%. This translates into a sharp decline for financial institutions other than pension funds, from 44.3 to 39.0% of the total.

"Crossholding" is, of course, one of the features that most distinguished the postwar Japanese economic structure, on par in importance with the main bank system, *keiretsu,* lifetime employment, and company-specific trade unions. Several merits have been ascribed to this system.

1) Corporate governance perspectives

The more stable shareholders a company has, the less risk there is that it will be the subject of a hostile takeover. Managers, who have usually been promoted from employees, are also able to run the company from a long-term perspective that emphasises the interests of the employees.

2) *Policy investment perspectives*

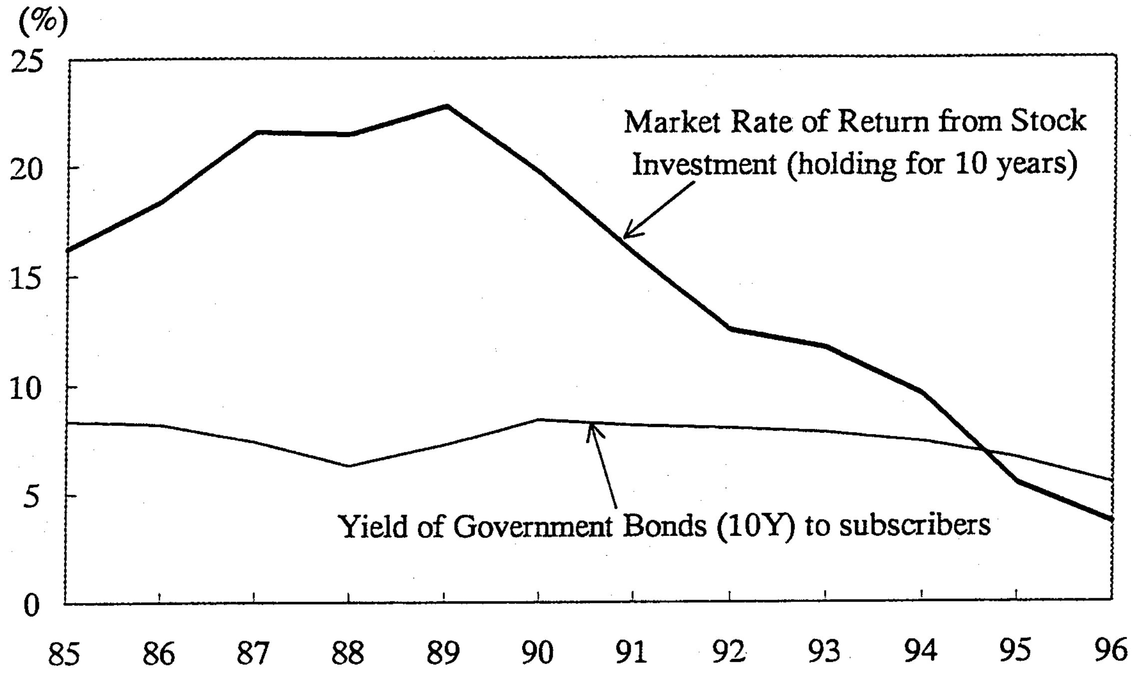
Crossholding enables companies to build long-term, stable trading relationships, which both reduces transaction costs and facilitates risk sharing. Shareholding arrangements between industrial companies and financial institutions lead to a reduction in "agency costs"; the institution is able to monitor corporate behaviour which reduces the credit risks, while the company is able to reduce its borrowing costs and increase the availability of loans.

3) *Higher unrealised profits*

The general rising trend for share prices gave shares in crossholding arrangements large unrealised profits that managers could use as a risk buffer. In other words, should the company be hit with an extraordinary loss that was difficult to cover out of recurring profits, it could realise the latent profits in its portfolio by selling shares at market prices and then, to re-establish the long-term relationships in its transactions, buying them back later on.

Figure 11

Market rate of return from stock and Government bond investment



Notes: The 1996 market rate of return on 10-year stock investment is the return/investment ratio when buying stocks at the average price in 1986 and selling them at the average price in 1996. Corresponding to market rate of return from stock investment, we used the average yield of 10-year Government bonds to subscribers in 1986 as the market rate of return from Government bond investment in 1996. Market rate of return on stock investment is calculated as follows (not only dividend but capital gain is added to the return from stock investment);

$$\textit{Market rate of return from stock investment} = \frac{\textit{Dividend} + \textit{Capital gain}}{\textit{Investment}} \times 100$$

$$= \frac{\textit{Dividend} + (\textit{Selling price} - \textit{Buying price})}{\textit{Buying price}} \times 100$$

Figures are calculated with a weighted average based on the aggregate market value of stocks listed on the First Sections of the Tokyo Stock Exchange.

Source: Japan Securities Research Institute, *Market Rate of Return from Stock Investment*.

The burst of the bubble in the nineties has changed this. In some cases, shares in crossholding arrangements have produced unrealised losses. In other cases, companies have had other losses to cover or needed to improve their cash flow and have, therefore, been forced to sell off crossheld shares for which there were still profits to be taken. From a macroeconomic perspective as well, expected growth rates have been in decline, but companies have needed to improve their earnings and meet the "structural adjustment" pressures, brought to bear by more intense international competition. This is forcing many to re-examine their business and capital relationships in the name of greater efficiency. We would point out that the prolonged share price slump has caused a substantial decline in the market average rate of return on equity investments (dividends plus capital gains divided by amount invested). Recently, equity investments held for a ten-year period have produced smaller returns than government bonds, which are considered a safe investment (Figure 11). These conditions will gradually force more and more companies to rethink their share crossholding arrangements, if only from the perspective of better investment efficiency.[19]

The internationalisation of investment yardsticks

In short, Japan is seeing its share crossholding arrangements unwind and a greater percentage of its shares going to foreign investors and domestic institutions (pension funds and the like), with signs of investment yardsticks moving in the direction of global standards. For example, there is a new emphasis on "return on equity" (ROE). The ROE of Japanese manufacturers has been in decline in the nineties because of the economic recession. Only recently has it bottomed out, but it is still not back to the average levels of the eighties, and the gaps with American companies are as wide as ever (Figure 12, top). Slumping ROE is basically a product of falling ROA (return on assets) (Figure 12, bottom). Improvements in ROE will require better investment efficiency and corrections to over-capitalisation. We would draw the reader's attention to the years 1984 and 1996, when there were roughly equal groupings of industries with rising and falling ROE. Compared with 1984, there were greater contrasts in the share prices' movements in 1996 (Figure 13). Obviously, there is no one single interpretation that can be put on these results. The economic environment was different in these two years and it is uncertain to what extent the markets had already discounted ROE in 1996, but it would be natural to see this as an indication that ROE was exerting a greater influence as an investment yardstick – not only were foreign investors emphasising ROE but domestic institutions have also been advocating greater use of ROE. These conditions are causing a greater number of corporate managers to explicitly list higher ROE among their business goals.

Another trend to be noted is the greater emphasis that institutional investors are putting on income gains, which has caused companies to compete on "payout ratios" and to make their dividends more elastic with respect to earnings levels. This represents an overhaul of traditional Japanese dividend policies, which were to minimise the amount of profit flowing out of the company and instead retaining profit inside for future investments, or to stabilise dividend amounts because crossholding relationships had produced a large contingent of stable shareholders. In the past, managers were content to let payout ratios swing widely over the business cycle.[20]

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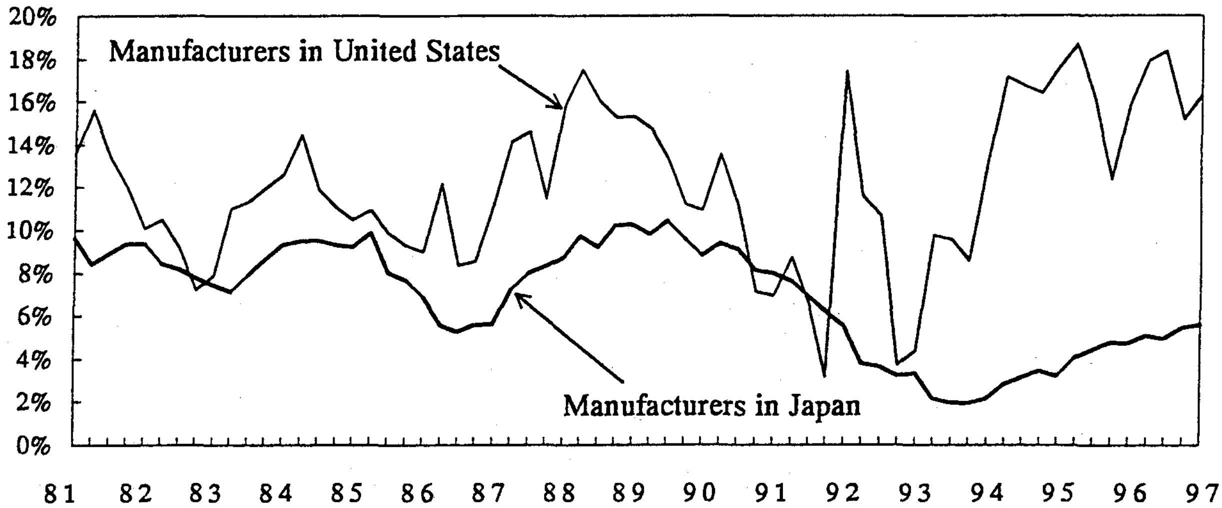
19 Nonetheless, it would be premature to think that crossholdings will immediately unwind. This is a practice that is deeply entwined with corporate governance and other aspects of the economic and corporate structure and is unlikely to disappear very rapidly or easily. Surveys indicate that many managers still see value in crossholdings. What will probably happen, therefore, is that crossholdings will be gradually unwound as managers become more selective about whose shares they hold.

20 As an illustration of the swings, the pay out ratio for all listed companies in Japan (2,267, including those in finance) was 30.3% in 1990, compared to 82.9% in 1994 and 60.8% in 1996.

Figure 12

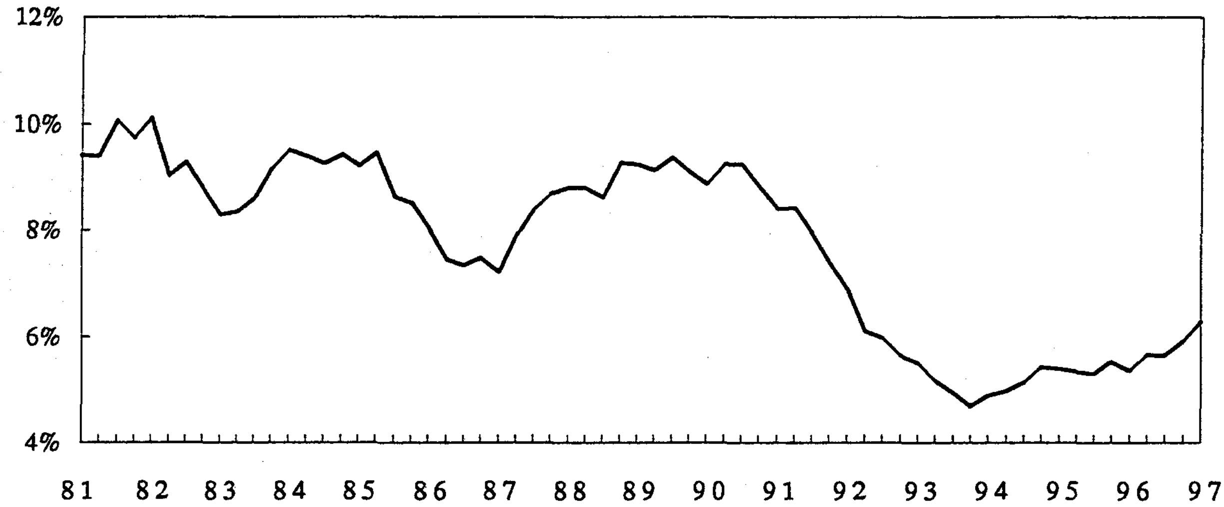
Rate of return on equity (Japan and the United States)



Notes: For Japan: *ROE = Current profits* x (1 - *Tax rate) / Own capital*; for the United States: *ROE = Profit for the current term after tax / Own capital.*

Sources: For Japan, Ministry of Finance, *Financial Statements of Incorporated Business, Annually and Quarterly*; for the United States, Department of Commerce, *Quarterly Financial Report for Manufacturing, Mining and Trade Corporations.*

Rate of return on assets (Japan)



Notes: *ROA* = (*Operating profit* + *Non-operating profit*) / *Assets*. The series is seasonally adjusted.

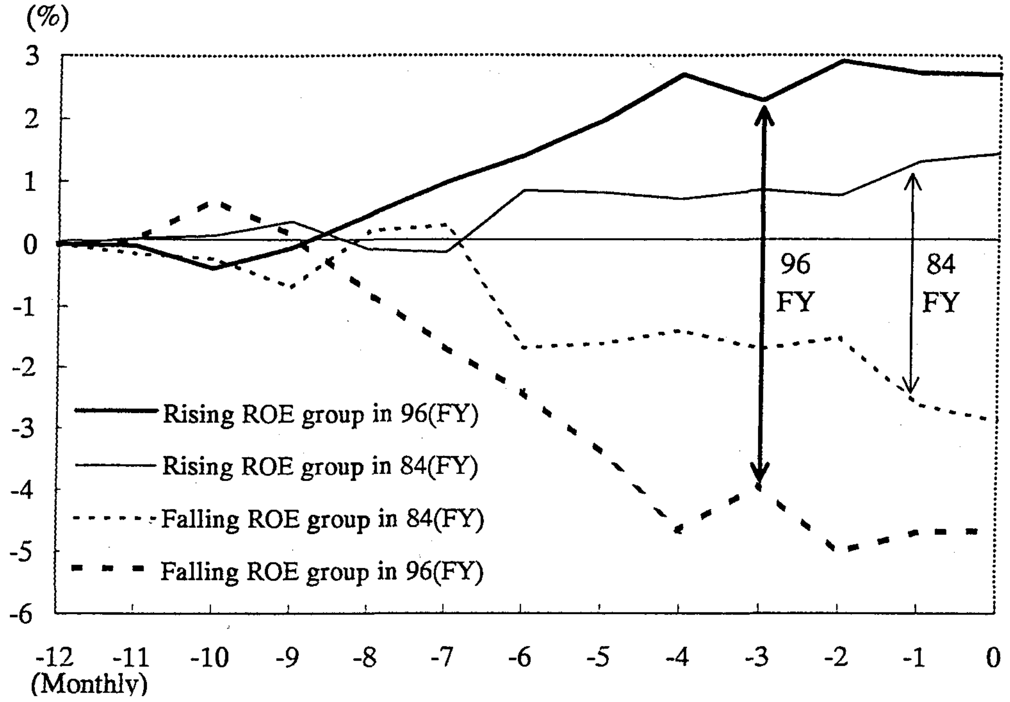
Source: Ministry of Finance, *Financial Statements of Incorporated Business, Quarterly.*

It is unclear to what extent these structural changes have really become established in the stock market. What we would point out to policy makers, however, is that changes are taking place. Hence, when they attempt to use share prices as an information variable, past experiences with the market may not always be reliable.

Figure 13

Changes in investment yardsticks (a new emphasis on ROE)

Contrasts in share price movements between groups of rising and falling ROE



Notes: We classified the 30 industries (excluding banking, insurance, securities) into two groups, based on rising or falling ROE. We averaged share prices of each groups to compare with the average of all industries, and looked at the contrast between share prices in rising and falling ROE by plotting share prices of each groups from 12 months before the publication of ROE. We selected 1984 (FY) and 1996 (FY) as samples because there were roughly equal number of industries in each grouping. 1984: rising = 20 industries, falling = 10 industries; 1996: rising = 19 industries, falling = 11 industries.

Sources: Tokyo Stock Exchange; Nihon Keizai Shimbun, Inc., *NEEDS (Nikkei Economic Electronic Databank System).*

Conclusion

This paper has so far verified the usefulness of share prices as information variables for policy makers and discussed the distinguishing characteristics of Japanese share price formation and the factors behind the slump of the nineties, particularly the role played by land prices. We have also touched on what appears to be signs of structural changes within the stock market during the nineties, emphasising the unwinding of crossholding relationships.

During the past year, the Nikkei average dropped from a high of 21,556 points[21] at the end of September 1996. During the January-March 1997 period it was hovering in the 17,000-18,000 point range. It later recovered to about 20,000 points during the May-July period, but has been slack since August. As of this writing in mid-September it was in the mid-17,000 point range. The major factors pushing share prices down during this period were uncertainties over the economic outlook caused by the fiscal austerity programme and a spate of corporate bankruptcies emerging in the

[21] This represents nearly a peak for post-bubble share prices. The Nikkei bottomed at 14,309 points in August 1992. Subsequent annual averages have been 19,100 for 1993, 19,935 for 1994, 17,329 for 1995, and 21,088 for 1996.

aftermath of the bubble. The low yield spread would indicate that there is little room for considering Japanese shares to be overvalued at current levels, but that does not mean the uncertainties over share prices will be resolved any time soon. Our observations so far in this paper indicate that three things will be required before share prices are able to begin a full-fledged recovery:

1) Recovery in the expected macroeconomic growth rate;

2) relief from the high credit risks brought by falling land prices – a cleanup of the negative legacies from the bubble;[22] and

(3) corporate behaviour emphasising shareholders values, such as the revision of dividend policies and improvement of return on equity (for example, by buying back shares from the market).[23]

Additionally, steps should be taken to introduce market valuation of assets and enhance disclosure requirements. During the boom and bust of the bubble, there were vast differences between the book values of assets on corporate accounts and their actual market values, and this made it difficult for investors to understand the assets and financial position of the companies they were investing in, increasing the opaqueness of investments. Other than these changes in corporate accounting, Japan also needs to improve its market infrastructure, for example, by establishing market practices that are both fair and transparent, reconsidering its securities taxation, and using deregulation to promote competition in the financial services sector. These realisations have inspired the government to move forward with a series of financial reforms, dubbed the "Japanese Big Bang". There are also structural reform plans for areas other than finance, and if the markets agree that the reforms will be effective, the consequent recovery in the expected growth rate should eventually be reflected in share prices.

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[22] Recent land price movements in urban areas indicate that considerable progress has been made at the macro-level in terms of the corrections required by the rupture of the bubble. Residential land appears to have stopped falling and commercial land prices are polarising between levels which are holding steady and levels which continue to drop, depending on the land's profitability. Overall, therefore, the rate of decline is contracting (of course, there are large differences among individual companies, including financial institutions, in the extent to which they have corrected their balance sheets).

[23] Until now, companies have rarely bought back their shares because of the "assumed dividends tax". This system was frozen in 1995, albeit only for three years and this combined with amendments to the Commercial Code in 1994 to cause a gradual increase in share buybacks. The amendments allow companies to buy their own shares if their shareholders agree to a profit write-off or if shares are needed to provide employees with stock options. Since 1995, sixty-three listed companies (Tokyo, Osaka, Nagoya) have bought back shares or have announced their intention to do so (as of 13th September 1997).

Appendix A: Correction for crossholding factors

The price/earnings ratio for the market as a whole is found by dividing market capitalisation by total earnings. The price/earnings ratio corrected for crossholding factors deducts cross-held shares from both market capitalisation and total earnings.

$$Adjusted\ market\ capitalisation = Apparent\ market\ capitalisation \times (1 - Crossholding\ ratio)$$

$$Adjusted\ total\ earnings = Apparent\ total\ earnings - Total\ dividends\ receivable\ from\ crossheld\ shares$$

$$= Apparent\ total\ earnings - (Total\ dividends \times Crossholding\ ratio)$$

$$= Apparent\ total\ earnings - (Apparent\ total\ earnings \times Payout\ ratio \times Crossholding\ ratio)$$

$$= Apparent\ total\ earnings \times (1 - Payout\ ratio \times Crossholding\ ratio)$$

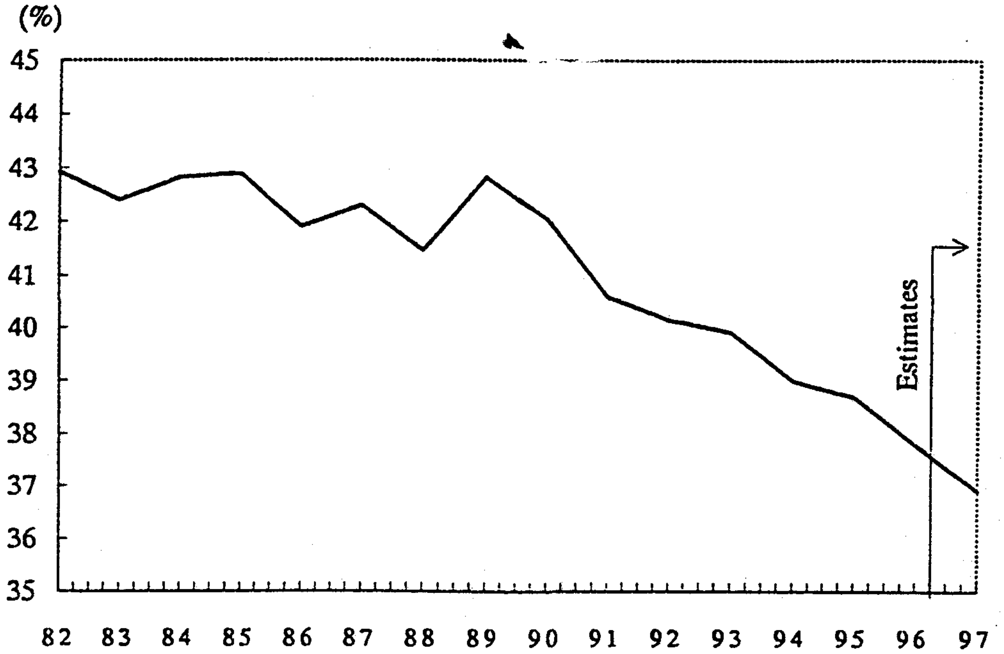
$$Adjusted\ price\ earnings\ ratio = \frac{Adjusted\ market\ capitalisation}{Adjusted\ total\ earnings}$$

$$= \frac{Apparent\ market\ capitalisation \times (1 - Crossholding\ ratio)}{Apparent\ total\ earnings \times (1 - Payout\ ratio \times Crossholding\ ratio)}$$

$$= \frac{1 - Crossholding\ ratio}{1 - Crossholding\ ratio \times Payout\ ratio} \times Apparent\ price\ earnings\ ratio$$

(Estimates by Daiwa Research Institute used for the crossholding ratio.)

Share crossholding ratio



Notes: Crossholding ratio of listed companies = Ordinary bank shareholding ratio + Trust bank shareholding ratio + Casualty insurance company shareholding ratio + Securities company shareholding ratio - Investment trust shareholding ratio - Pension trust shareholding ratio - Public fund shareholding ratio - Tokkin and fund trust shareholding ratio + Other corporate shareholding ratio * 0.7.

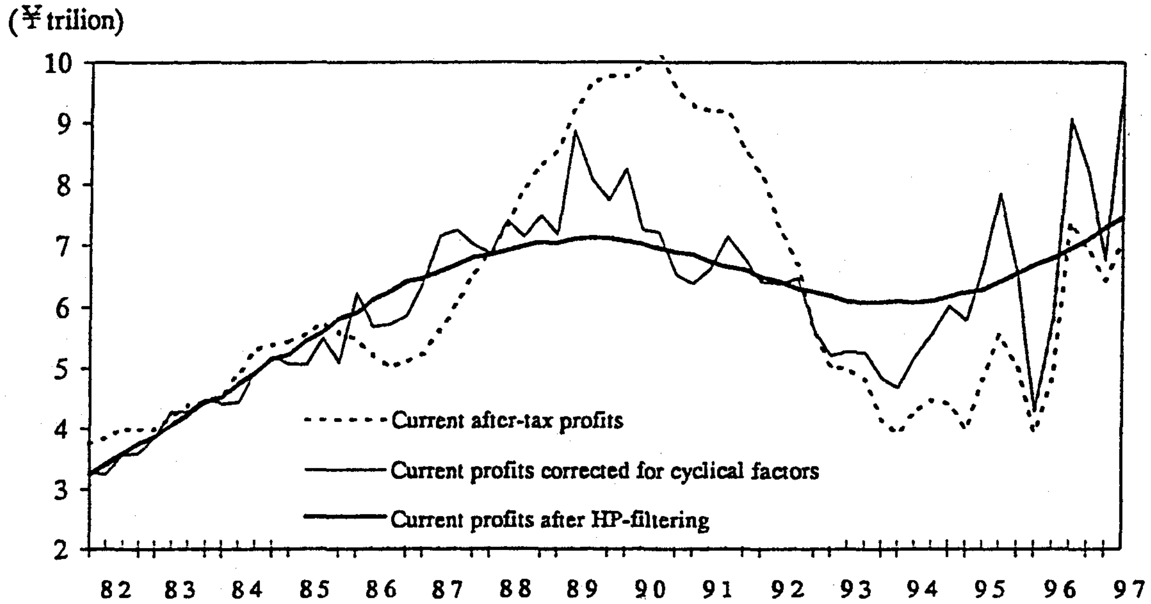
Estimates for the second quarter 1996 and beyond assume that unwinding proceeded at the same pace as during the 1995-96 fiscal years. First quarter figures for each year are from Daiwa Research Institute (other quarterly figures were as indicated by the graph lines).

Appendix B: Correction for cyclical factors

Short-term corporate earnings will undergo large swings because of business cycles, and the earnings that the markets use to forecast the future stream of corporate earnings are based on the assumption that the expected growth rate of nominal earnings is constant, and may differ from actual earnings. In other words, if the economy is currently in recession but corporate earning are forecast to recover in the future, then the price/earnings ratio will be upward biased, while if the economy is currently in a boom but corporate earnings are forecast to decline, the opposite will be true. Therefore, when assessing price/earnings ratios, it is necessary to eliminate these cyclical factors from calculations of corporate earnings.

There are many techniques that could be used to correct for cyclical factors. The technique we have used is to take the residual from a regression of forecast earnings on the GDP gap (estimated), and to assume that there is a trend after elimination of cyclical factors. We then use the residual from the previous calculation and substitute the average gap value during the estimation period for the gap effect, thereby arriving at a forecast earnings trend corrected for cyclical factors. To this we apply an HP filter (l = 1,600) to smooth out the curve and eliminate noise. These values have been used in this paper as "corporate earnings corrected for cyclical factors".

Correction of corporate earnings for cyclical factors



Estimation formulas:

(1) *LN* ((Current after-tax profits (real)) = 16.04 + 0.12 x *GDP gap*
(221.3) (6.5) ( ) = t-value

Estimation period = 1982Q1~1997Q1

Adjusted R-square = 0.404 S.E. = 0.226 D.W. = 0.340

(2) *Current after-tax profits* (real; corrected for cyclical factors)

= EXP (16.04 + 0.12 x *Average value for GDP gap during the estimation period*) $+\varepsilon$, where ε is the residual from Equation (1).

(3) An HP-filter (l = 1,600) is applied to the values from Equation (2), and the results deemed current after-tax corporate profits corrected for cyclical factors.

Notes: The seasonally adjusted GDP deflator was used to compute real values. The graph shows nominal current after-tax profits.

Asset prices and monetary policy in Sweden

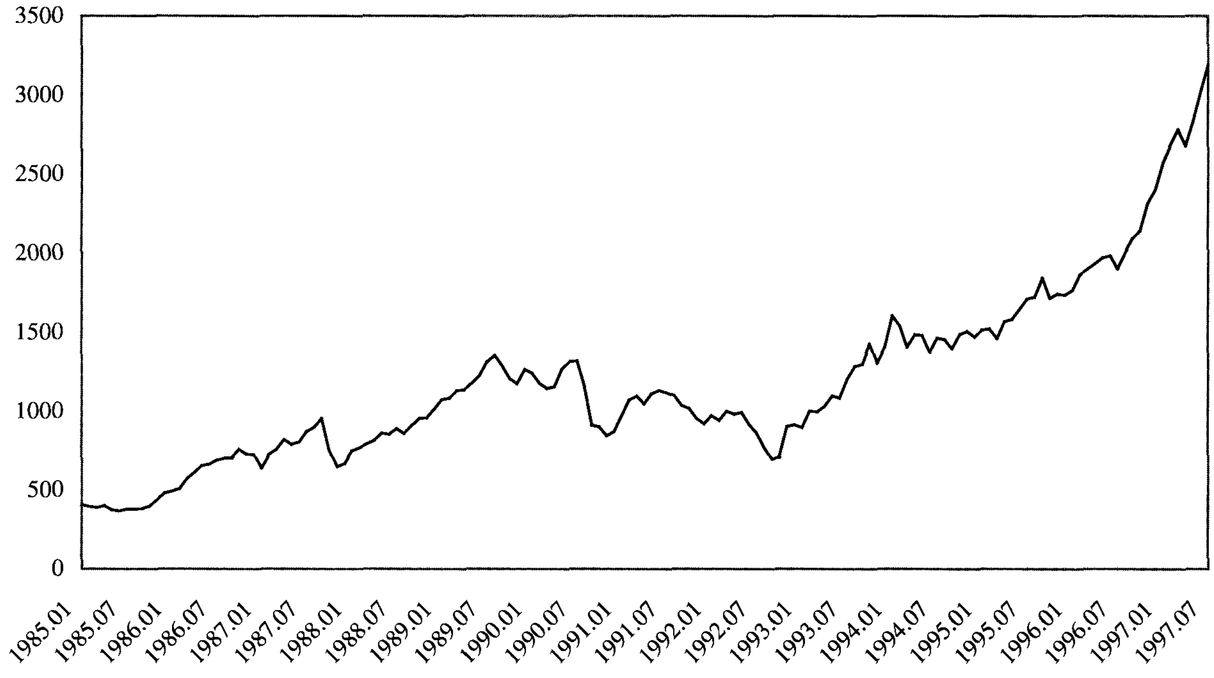
Peter Sellin

Introduction

Up until 7th August 1997 the Swedish stock market had risen without a major correction since Sweden abandoned the fixed exchange rate system in November 1992. The recent correction has been around 8% as in most European stock markets. Figure 1 shows the Affärsvärlden's General Index (AFGX) for the period January 1985 to July 1997. During this period there have been three major setbacks. The first was in 1987 when the market fell by 32% in just two months. The second major reversal in stock prices came in 1990 when the market dropped by 36% between July and November. The third correction occurred in 1992 with a 30% decrease from May to September.

Figure 1

AFGX monthly index 1985 to 1997



Looking at this recent history it is only natural to wonder when (not if) the next major correction will occur. However, this may not be correct. It could be that the 1980's is an exceptional period and that today's stock prices actually reflect strong underlying fundamentals. We address this question in Section 1 by adopting a longer perspective to model the underlying fundamentals in a simple real asset-pricing framework. Substantial deviations from the fundamental price are found. However, it is recognised that the deviations from fundamentals could be related to monetary policy. In Section 2 we consider a model where monetary policy has real effects, and Section 3 investigates whether various policy actions by Sveriges Riksbank have had any impact on the stock market.

1. Do equity prices reflect fundamental values?[1]

We use a simple exchange economy of the Lucas (1978) type with a representative agent with constant relative risk aversion utility. Assuming a (real) dividend process of the form

$$D_{t+1} = D_t e^{\alpha + \varepsilon_{t+1}} \tag{1}$$

where $\alpha > 0$ and $\varepsilon_t \approx N(0, \sigma^2)$, this model can be shown to have as solution the following fundamental price,

$$P_t^* = \rho D_t \tag{2}$$

where ρ is a constant (and a function of parameters of the dividend process and the utility function). Hence, we get a very convenient solution, with the fundamental price as linear function of today's dividend.

We use an annual index of Swedish stock returns constructed by Frennberg and Hansson (1992a, 1992b) and later updated. This is a value-weighted index that includes dividends. It is constructed along the same principles as in the standard work by Ibbotson and Sinquefield (1989) and spans the period from 1919 to 1996. A dividend series and an index of consumer prices, which are also part of the Frennberg-Hansson data set, were also used, with the latter applied to compute real returns and real dividends.

Using the average price-dividend ratio for the whole sample as a proxy for ρ we can derive a fundamental price series. This series is plotted along with the actual price series in Figure 2. The corresponding series in real terms are shown in Figure 3. Three distinct subperiods can be distinguished. Up until a few years after the second world war the actual price was consistently below the fundamental price. In the post-war period up until the early 1980s the actual price fluctuated around the fundamental. In the 1980s and 1990s the actual price has been above the fundamental price. There have been three partial collapses of the price bubble during this latter period. The third of these almost brought the price back to its fundamental value. During 1992-93 the two price series drastically part company when the fundamental price drops significantly while the actual price rises. During the two previous episodes, when the fundamental price dropped dramatically in the early 1920s and early 1930s, the actual price followed the fundamental. This was not the case in 1992-93. The explanation for this is the following. Between the summers of 1990 and 1993, GDP dropped by a total of 6%, dealing a heavy blow to Swedish companies' earnings and to the fundamental price.

During the currency turmoil in the autumn of 1992, Sweden had to abandon the fixed exchange rate on 19th November. This resulted in an immediate de facto devaluation of 12% against the dollar.[2] This was, of course, expected to lead to an improved competitive situation for the export-oriented Swedish manufacturing industry. This is one reason for the rising equity prices after November 1992. Another reason is that Sveriges Riksbank started lowering interest rates. A third reason is the abandonment of restrictions on foreign ownership of Swedish equity on 1st January 1993.[3]

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1 This section draws on Nydahl and Sellin (1997). See also Sellin (1997a) for a similar analysis of the US stock market.

2 Close on 20th November compared to close on 19th November.

3 See Sellin (1996) on the effects of lifting these exchange controls.

Figure 2

Actual and fundamental equity prices

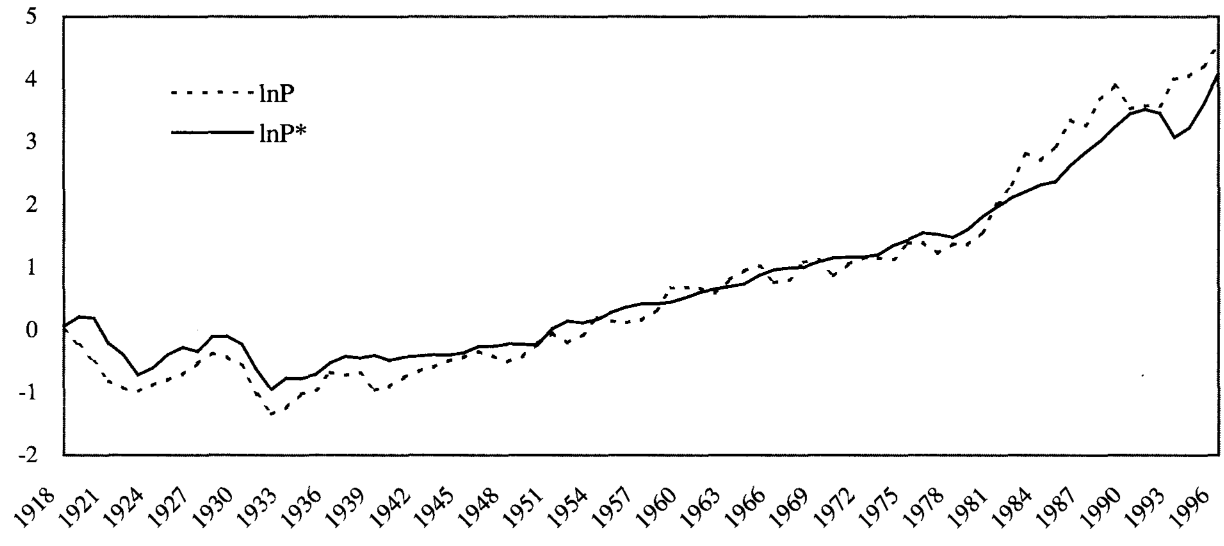
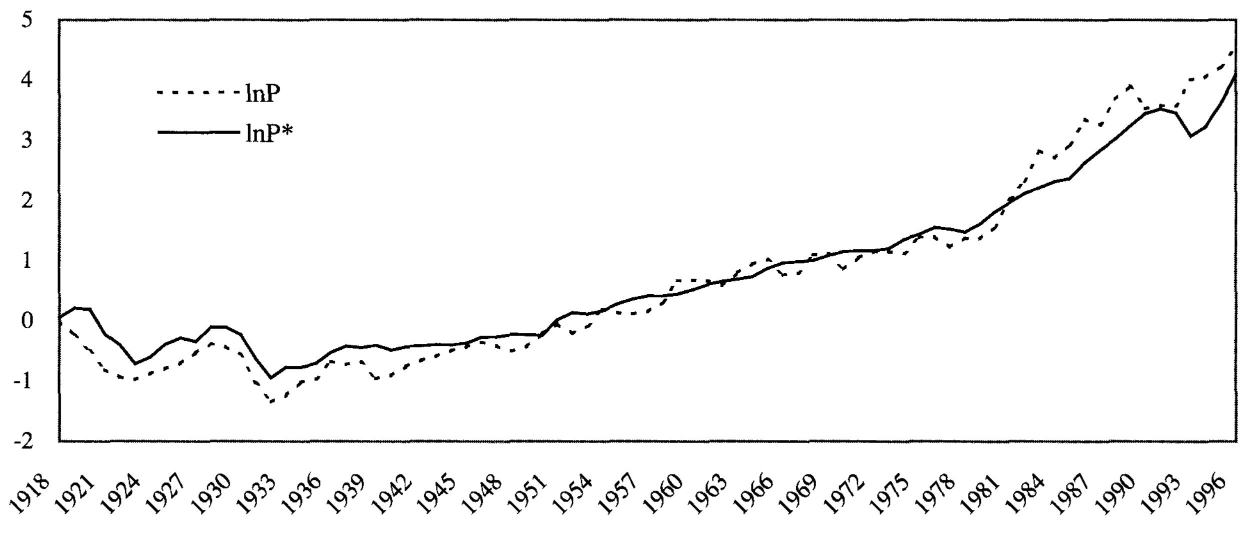


Figure 3

Real actual and fundamental equity prices



Towards the end of the sample period there is a strong rebound in the fundamental price but not a closing of the gap, since the actual price also rises quite significantly during this period. However, it looks as if the fundamentals are on their way to catching up with the expectations driving the actual price.

It seems unlikely that the stock market has been undervalued from 1919 to 1945. It is more likely that we have overestimated the fundamental price up until 1945. The gap at the very end of the sample could likewise be due to the fundamental price being underestimated. A higher growth rate or variance in the dividend process would imply a higher ρ and thereby a higher fundamental price. Our model uses the same ρ for the whole sample period. These issues are discussed more fully in Nydahl and Sellin (1997), where the model is also estimated. In that paper, to formally test for the existence of bubbles we adapt a switching regime approach suggested by van Norden and Schaller (1994). The results are somewhat mixed and most of the testable implications from our theoretical model fail to find significant support in the data. There is some evidence of a bubble in equity prices in the 1980s. From Figure 2 it looks as if the bubble economy has continued into the 1990s. However, in a model where monetary policy has real effects a different interpretation could be given to the deviations from the "fundamental price". We examine this possibility in the next section.

2. Nominal asset pricing models[4]

Introducing money into a general equilibrium asset-pricing model is not a trivial undertaking. We will follow Lucas (1980, 1982) and require that the agent has to meet a cash-in-advance constraint for purchasing the consumption good.

The model is set up in the following way. The representative agent enters a period with money and equity shares carried over from the previous period. He receives a helicopter drop of money and the securities market opens for trading. The security market closes and the goods market opens for trading. Goods must be bought with money (currency). The goods market closes and the agent collects dividends in the form of currency, which is carried into the next period.

The agent's problem is to choose consumption, c, money holdings, M, and equity shares, z, given the price of the good, p, and the real price of equity, q, so as to maximise:

$$E_0 \sum_{t=0}^{\infty} \beta^t u(c_t)$$

subject to a budget constraint,

$$\frac{M_t}{p_t} + q_t z_t \leq \left(q_t + y_{t-1} \frac{p_{t-1}}{p_t} \right) z_{t-1} + \frac{M_{t-1} - p_{t-1} c_{t-1}}{p_t} + \frac{M_{t+1}^s - M_t^s}{p_t}, \qquad t \geq 0,$$

and a cash-in-advance constraint,

$$M_t \geq p_t\, c_t \qquad t \geq 0 .$$

The restrictions will be binding at the optimum solution. In equilibrium we also require that for every $t \geq 0$:

$$c_t = y_t, \; z_t = 1, \text{ and } M_t = M_{t+1}^s .$$

Using the binding cash-in-advance constraint and substituting it and the equilibrium conditions into the binding budget constraint, we can derive a theoretically determined price level:

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4 This section draws on Sellin (1997b).

$p_t = M^s_{t+1} / y_t$.

Substituting this price level and the equilibrium condition for consumption into the Euler equation for the equity price, we end up with the equilibrium price of equity:

$$q_t = \beta E_t\left[\frac{u'(y_{t+1})}{u'(y_t)}(q_{t+1} + y_{t+1}/\mu_{t+1})\right],$$

where $\mu_t = M^s_{t+1} / M^s_t$ is the growth rate of money. Assuming that the money supply process is independent of the endowment process, we get

$$q_t = \beta E_t\left[\frac{u'(y_{t+1})}{u'(y_t)}q_{t+1}\right] + \beta E_t\left[\frac{u'(y_{t+1})}{u'(y_t)}y_{t+1}\right]E_t\left[\frac{1}{\mu_{t+1}}\right].$$

From this equation it is clear that the effect of increased expectations of a (temporary) monetary tightening will have a positive effect on the real equity price,

$$\frac{\partial q}{\partial E_t\left[\frac{1}{\mu_{t+1}}\right]} > 0.$$

A monetary tightening is expected to lead to lower inflation and a higher purchasing power of the dividend sum carried over to the next period.

Boyle (1990) gets the opposite result to the one derived above, for an agent with low constant relative risk aversion.[5] For an agent with higher risk aversion the effect is ambiguous in Boyle's model. He uses a money-in-the-utility-function model with variable velocity of money. Marshall (1992) derives a similar result in a model where money economises on transactions costs. The intuition is the same in the two models. Expectations of monetary easing leads the agent to substitute out of money and into equities, thus raising the real price of equity. Hence, whether expectations of a monetary tightening/easing has a positive or negative effect on real equity prices is an empirical question. We turn to this in the next section.

3. The impact of Swedish monetary policy on the stock market

There have been a number of studies of the impact of monetary policy on asset prices. Most of these look at the ability of monetary policy to influence money market interest rates.[6] The earlier literature is reviewed in Reichenstein (1987). More recent studies have been made using US data (Cook and Hahn (1989), Tarhan (1995)), UK data (Dale(1993)), and data for the G10 countries (BIS (1997)). However, there are few studies that have considered the impact of monetary policy on equity prices. Tarhan (1995) considers the impact of Federal Reserve open market operations on financial assets other than interest rates (in a study mainly focusing on interest rates). He finds no evidence that the Fed influences stock prices. Thorbecke (1997), on the other hand, finds a significant negative effect on the percentage change in the Dow Jones Industrial Average from policy-induced

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5 A CRRA parameter of less than one.

6 For the Swedish case see Lindberg, Mitlid and Sellin (1997).

changes in the federal funds rate. The different choice of policy instrument in these two studies follow naturally from the choice of sample period.[7]

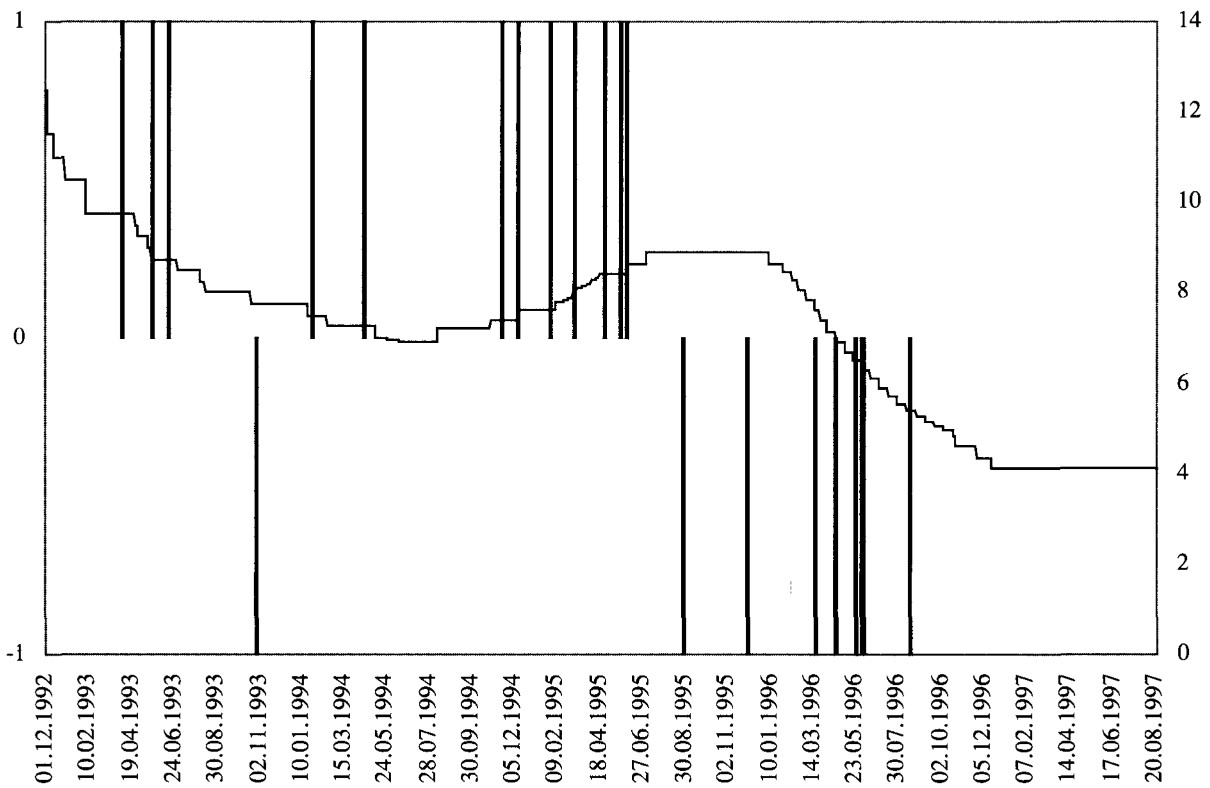
In this section we will focus on the impact of the Sveriges Riksbank policy instruments on the stock market. We will consider a wider set of instruments than have been used in any previous study. We start with a description of these instruments.

The Swedish system for the practical management of monetary policy was introduced in June 1994. It provides one deposit and one lending facility. The deposit and lending rate are set by the Governing Board of the Riksbank and form a corridor within which the repo rate – the Riksbank's primary instrumental rate – is set by the Governor in accordance with monetary policy guidelines established by the Governing Board. The interest rate corridor provides the Riksbank with a tool for signalling its long-term intentions concerning the repo rate.

The repo rate is the rate at which, as a means of managing the liquidity of the banking system, securities with a maturity of one week are bought or sold by the Riksbank under a repurchase agreement. The repo rate may be interpreted as the Riksbank's target for the level of the overnight rate in the interbank market. Repos or reversed repos are placed by tender every Tuesday. Repos are normally offered at a fixed rate, leaving the Riksbank's counterparties to tender the volumes they are interested in depositing or borrowing for one week at that rate.

Figure 4

Speeches and the repo rate



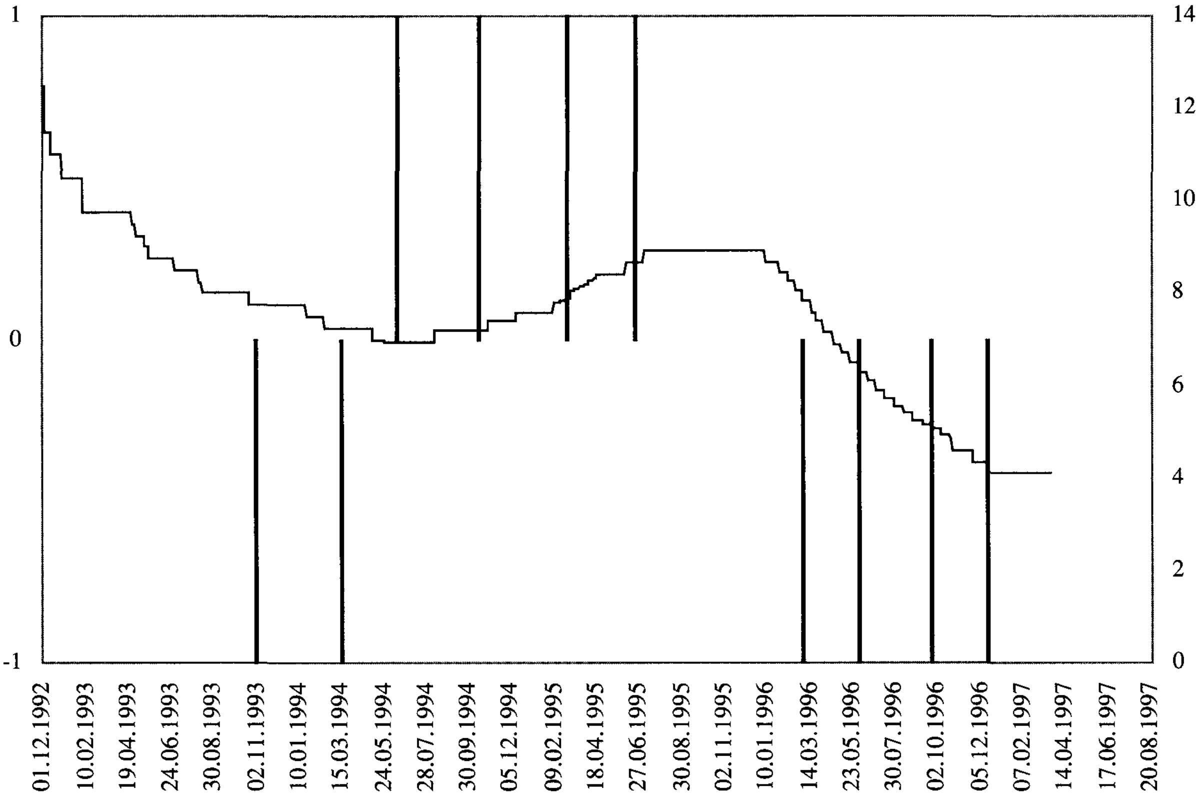
7 Tarhan studies the period 2nd October 1979 to 31st December 1984 when the growth rate of money was the target. Thorbecke focuses attention on periods when the federal funds rate was targeted. He uses Cook and Hahn's (1989) 1974-79 fed funds data and adds a similarly constructed series for the period 11th August 1987 to 31st December 1994.

The Riksbank's intention has been to be transparent in its monetary policy considerations since the explicit inflation target was introduced in January 1993. There are various ways in which a central bank can influence expectations about monetary policy. The traditional channel of information is speeches and lectures by the Governor and staff of the Riksbank. The Riksbank also issues an inflation report four times a year to present its assessment of future inflation and the implications for monetary policy to the financial markets and to the public. In this way, the markets get an indication of the Riksbank's intentions and changes in monetary policy will not come as a surprise.

We will be looking at any potential impact on the stock market from any of these monetary policy instruments. In order to assess effects from inflation reports and speeches by the governor and deputy governors of the Riksbank, these have been coded 1(-1) if the report/speech was interpreted (by the author) as foreboding monetary tightening (easing) and the value zero if the report/speech was neutral. These dummy variables are shown along with the repo rate in Figures 4-5.[8] During the first phase of monetary easing the speeches seem to have served mostly as warnings that the lowering of the repo rate may not proceed at the pace expected by the market. During the next two phases the speeches seem to have served rather to prepare the market for coming repo rate changes. The inflation reports in Figure 5 have also been in line with subsequent changes in the repo rate.

Figure 5

Inflation reports and the repo rate



When investigating whether monetary policy has an impact on the stock market, it may be interesting to consider the impact on both returns and volatility. We can do this simultaneously by

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8 The dates for the inflation reports and for announcing changes in the repo and lending rates can be found in the Appendix.

using a model of the ARCH family (see Engle, Chou and Kroner (1992) for an overview of these types of models). Not only the conditional mean but also the conditional variance, h_t, is modelled. The latter is modelled as a function of lagged squared residuals, u^2_{t-1}. It is also possible to include exogenous or predetermined variables in both the mean and variance equations. We will include our policy variables in the mean equation and the absolute values of the policy variables in the variance equation.

Table 1

Policy impact on stock and bond markets

Variable	R_t^{AFGX}		Δr_t^{SE}	
Mean equation intercept	0.081	(3.198)	-0.006	(2.557)
Dependent variable (*t*-1)	0.020	(0.731)	0.058	(1.791)
Dependent variable (*t*-5)			-0.099	(3.469)
R_{t-1}^{SP500}	0.420	(11.500)		
Δr_{t-1}^{GE}			0.091	(1.885)
$\Delta REPO_t$	-1.195	(2.226)	0.081	(2.209)
$\Delta REPO_{t-1}$			0.072	(1.471)
$\Delta LEND_t$	0.197	(0.553)	0.055	(1.322)
$SPEECH_t$	0.081	(0.533)	0.018	(1.143)
$REPORT_t$	-0.340	(1.118)	-0.010	(0.328)
Variance equation intercept	0.042	(2.448)	0.001	(4.948)
u_{t-1}^2	0.055	(2.485)	0.201	(7.909)
h_{t-1}	0.890	(23.270)	0.700	(22.220)
$\lvert\Delta REPO_t\rvert$	0.569	(1.668)	0.002	(0.457)
$\lvert\Delta LEND_t\rvert$	-0.385	(1.744)	0.010	(1.804)
$\lvert SPEECH_t\rvert$	-0.216	(2.979)	0.002	(1.273)
$\lvert REPORT_t\rvert$	0.177	(0.918)	0.001	(0.516)
Ljung-Box Q(10)	11.719	[0.3043]	9.744	[0.4632]
Ljung-Box Q*(10)	4.618	[0.9152]	9.388	[0.4958]
Bera-Jarque	0.477	[0.7877]	164.5	[0.0000]

Notes: t-values are reported within parentheses and probability values within square brackets. The variable *REPORT/SPEECH* is a dummy variable that takes the value 1 (-1) if the inflation report/speech was interpreted (by the author) as foreboding monetary tightening (easing) and the value 0 if the report/speech was neutral. Ljung-Box tests for autocorrelation in the standardised residuals Q and the squared standardised residuals Q* respectively. Bera-Jarque tests the assumption that the standardised residuals are normally distributed.

In Table 1 we report the results from estimating a GARCH(1,1) model for the daily return on the AFGX equity index, R_t^{AFGX}. As a comparison, the same type of model has also been estimated for the change in the five-year government bond yield, Δr_t^{SE}. The foreign influence has been considered by including the lagged return on the S&P 500 index in the AFGX model and the change in the German five-year government bond yield in the interest rate model. Both foreign influences are positive and significant as expected.[9] The Bera-Jarque statistic warns us that the standardised residuals from the interest rate model are not normally distributed so we have to be a bit careful with drawing strong conclusions with regard to the inference from this model.

We consider the effect on the conditional mean first. A change in the repo rate has the expected negative effect on equity returns and positive effect on the interest rate. There is no evidence of a lagged effect of the repo rate on stock returns, which is why the model was re-estimated without the lagged repo variable. The lending rate and speeches have the wrong sign in the AFGX model but are not statistically significant. The inflation report has the right sign but is not significant either.

Turning to the conditional variance, the GARCH parameters are highly significant and volatility displays the high persistence, usually found in daily data (the sum of the parameters is close to one). As expected, there is a positive effect on volatility from the announcement of a change in the repo rate (significant at the 10% level). There is a significant *negative* effect from changes in the lending rate and from speeches, while the inflation report has no effect. One way to interpret these results is that both changes in the lending rate and speeches by the governor and deputy governors remove uncertainty about the future path of monetary policy, which results in lower volatility in the market. This interpretation, of course, begs the question as to why the effect on volatility is not the same in the bond as in the stock market, though, as far as the speeches are concerned, only the effect on stock market volatility is significant (1% level).

Conclusions

The Swedish stock market has risen in value with no major correction since November 1992. Our question is if we should expect to see a major decline in the near future and whether prices still reflect fundamental values? In this paper, a measure of fundamental value was computed from a simple asset-pricing model and compared with the actual prices. The actual price since 1992 has been above the fundamental; but there has also been a strong increase in the fundamentals in the past few years and the gap could be closing.

A follow-up question is if the central bank can influence the stock market. The impact on the stock market from changes in the monetary policy instruments of Sveriges Riksbank was examined. It was found that Sveriges Riksbank indeed influences the level of equity prices as well as the volatility in the market. Whether it is desirable use these instruments to cool off the stock market, in view of the uncertainty regarding the difference between fundamental and actual values, is a different question.

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9 For an analysis of US and German interest rate and volatility transmission to the Swedish money and bond markets see Dahlquist, Hördahl and Sellin (1997).

Appendix: Changes in policy instruments, December 1992 to August 1997

No.	Announcement Date	Repo rate	Repo rate change	Lending rate change	Inflation report code
1	1992-12-03	11.50	-1.00		
2	1992-12-14	11.00	-0.50		
3	1992-01-05	10.50	-0.50		
4	1993-02-05	9.75	-0.75		
5	1993-04-23	9.50	-0.25		
6	1993-04-29	9.25	-0.25		
7	1993-05-13	9.00	-0.25		
8	1993-05-19	8.75	-0.25		
9	1993-07-01	8.50	-0.25		
10	1993-08-05	8.25	-0.25		
11	1993-08-12	8.00	-0.25		
12	1993-10-21	7.75	-0.25		
13	1993-10-29				-1
14	1994-01-20	7.50	-0.25		
15	1994-02-17	7.25	-0.25		
16	1994-03-14				-1
17	1994-05-05	7.00	-0.25		
18	1994-05-26	6.95	-0.05		
19	1994-06-14	6.92	-0.03		0
20	1994-08-11	7.20	0.28	0.50	
21	1994-10-18				1
22	1994-11-01	7.40	0.20		
23	1994-12-13	7.60	0.20		
24	1995-02-09	7.80	0.20	0.50	
25	1995-02-21	7.83	0.03		
26	1995-02-28	7.90	0.07		1
27	1995-03-07	8.05	0.15		
28	1995-03-14	8.10	0.05		
29	1995-03-21	8.15	0.05		
30	1995-03-28	8.20	0.05		
31	1995-04-04	8.27	0.07		
32	1995-04-11	8.34	0.07		
33	1995-04-12			0.50	
34	1995-04-18	8.41	0.07		
35	1995-06-06	8.66	0.25		
36	1995-06-20				1
37	1995-06-29			0.50	
38	1995-07-04	8.91	0.25		
39	1995-11-16				0
40	1996-01-09	8.66	-0.25		
41	1996-01-30	8.45	-0.21		
42	1996-02-13	8.30	-0.15		
43	1996-02-22	8.05	-0.25	-0.50	
44	1996-03-04				-1
45	1996-03-05	7.85	-0.20		
46	1996-03-19	7.60	-0.25		
47	1996-03-21			-0.75	
48	1996-03-26	7.40	-0.20		
49	1996-04-09	7.15	-0.25		

No.	Announcement Date	Repo rate	Repo rate change	Lending rate change	Inflation report code
50	1996-04-23	6.90	-0.25		
51	1996-04-25			-0.75	
52	1996-05-07	6.70	-0.20		
53	1996-05-21	6.50	-0.20		
54	1996-06-04	6.30	-0.20		
55	1996-06-05				-1
56	1996-06-18	6.10	-0.20		
57	1996-06-20			-0.75	
58	1996-07-02	5.90	-0.20		
59	1996-07-16	5.70	-0.20		
60	1996-07-30	5.55	-0.15		
61	1996-08-13	5.40	-0.15		
62	1996-08-15			-0.50	
63	1996-08-27	5.25	-0.15		
64	1996-09-10	5.15	-0.10		
65	1996-09-24	5.05	-0.10		-1
66	1996-10-08	4.95	-0.10		
67	1996-10-22	4.85	-0.10		
68	1996-10-24	4.60	-0.25	-0.50	
69	1996-11-26	4.35	-0.25		
70	1996-12-05			-0.50	
71	1996-12-17				-1
72	1996-12-18	4.10	-0.25		
73	1997-03-20				0
74	1997-06-05				0

Notes: The change in the deposit rate has been the same as for the lending rate except on two occasions. On 11th August 1994 the deposit rate was not changed and on 12th April 1995 it was raised by twice as much. The inflation report code takes the value 1 (-1) if the report was interpreted (by the author) as foreboding monetary tightening (easing) and the value zero if the report was neutral. The same procedure was used in coding the speeches by the governor and deputy governors of Sveriges Riksbank (not reported here).

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Equities: what can they tell us about the real economy?

Simon Hayes, Chris Salmon and Sanjay Yadav

Introduction

A feature of most industrialised countries in the recent past has been the strong growth in equity prices. This poses many questions to policymakers, chief amongst which are: what has led to the increase in equity prices, and what are the implications of significantly higher equity prices for the rest of the economy? This article draws together several disparate strands of research that attempt to address these issues and that are on-going at the Bank of England.

The next two sections focus on explanations of the increase in UK equity valuations. We first discuss the equity risk premium, with a view to finding out whether it has fallen in recent years compared to its long-run average. The main alternative explanation for higher equity prices is that expectations of future dividend growth have increased, and we discuss evidence relating to this hypothesis in Section 2.[1]

Thereafter we focus on the possible implications of the rise in equity prices. Monetary authorities may care about developments in equity prices for a variety of reasons. At the simplest level, equities may act as leading indicators for developments elsewhere in the economy. A priori this is a plausible supposition, given that a fundamental determinant of equity prices is expected future corporate earnings. An increase in equity prices, for example, driven by an upwards re-assessment of future corporate earnings might provide early evidence of a positive demand or supply shock. Or more structurally, as discussed in Section 3, changes in equity prices may themselves form part of the transmission mechanism of monetary policy. Changes in equity prices will change the net worth of both consumers and corporates, and such changes may have additional direct effects upon both consumption and investment, over and above those arising from the change in the cost of capital.

The maintained assumption throughout these three sections is that equity prices reflect fundamentals. We do not consider the possibility and implications of price bubbles, but focus on "no bubbles" analysis, that we think in general more instructive.[2]

In Section 4 we present some preliminary analysis of the leading indicator properties of equity prices in the United Kingdom, which is designed to clarify the informational content of equity market. The results in this section are largely incomplete however, and we discuss how we intend to extend this analysis.

Finally, in Section 5 we switch attention to stock option prices. We discuss how option prices can be utilised to extract the implied distribution of expected future stock prices at option

1 The other possible explanation is a change in the discount rate which the market applies in valuing expected future earnings, and the discount rate could be affected directly by official interest rate changes.

2 There are a number of reasons, theoretical and empirical, to discount the likelihood of rational bubbles. From a theoretical angle, a negative bubble can never exist on an asset with limited liability, and a positive bubble can only exist if investors believe that there is no upper bound on stock prices. The latter rules out the possibility that firms issue more equity once the price reaches a certain level, thus effectively capping the stock price. Also, a bubble can *never* be zero. This means that if a bubble exists it must have always existed; and if it goes to zero, it can never re-start. Finally, note that rational bubbles are not predictable, and so cannot be proffered as an explanation for predictability in stock returns. From an empirical perspective, rational bubbles imply explosive behaviour in functions of prices and dividends that are not consistent with observed behaviour (see Campbell et al. (1997)).

maturity and may, therefore, themselves contain leading indicator information for equity prices. The final section summarises.

1. Equity risk premium

Many analysts are of the opinion that a substantial portion of the rise in the UK stock market may be due to the fall in the equity risk premium. The equity risk premium is the additional return investors require as compensation for bearing the risks associated with holding equities, compared with risk-free assets. A lower-risk premium on equities implies that agents will use a lower discount rate or required rate of return to discount future dividend payouts; *ceteris paribus*, this should mean that the market rate will rise. Additionally, a lower equity risk premium will lead to a fall in the equity cost of capital which might then induce higher investment spending by the corporate sector (this is discussed in more detail in Section 3 below).

We model the ex-ante evolution of the risk premium in order to find out whether the current level of the equity risk premium is lower than the historical average. To that end, we use a dynamic version of the CAPM model developed by Merton (1973). Merton's model involves a market with continuous trading, where investors' utility falls as expected volatility (measured by the instantaneous conditional variance of asset returns) increases. In equilibrium, there is a linear relation between the required return on the market portfolio over and above the risk-free interest rate (i.e. the market risk premium), and the conditional variance of returns on the market portfolio:

$$E_t(R_{t+1} - r_{t+1}) = \gamma\sigma_{t+1}^2 \tag{1}$$

where R_{t+1} is the required return on the market portfolio in period $t+1$, r_{t+1} is the risk free rate in period $t+1$ and σ^2_{t+1} is the conditional variance of returns on the market portfolio. E_t is the expectation formed using information available at time t. The coefficient γ is commonly interpreted as a measure of average risk aversion.

To implement equation (1), we need some measure of the expected market return variance. Following Nelson (1991) we use an EGARCH-M specification to model the conditional variance of excess returns:[3]

$$R_{t+1} - r_{t+1} = \gamma\sigma_{t+1}^2 + \varepsilon_{t+1} \tag{2}$$

$$\varepsilon_{t+1} = \sigma_{t+1} z_{t+1} \qquad z_{t+1} \sim iid(0,1)$$

$$\ln\sigma_{t+1}^2 = \alpha_0 + \alpha_1 \ln\sigma_t^2 + \theta z_t + \gamma(|z_t| - E|z_t|) \tag{3}$$

Since EGARCH models the *log* of the return variance, rather than the level, the variance will be positive regardless of the sign of the estimated parameters. A particularly attractive feature of EGARCH is that it allows for asymmetry in the response of the conditional variance to positive and negative shocks to returns. Assuming that $\gamma > 0$ (as it is usually found to be), if $\theta < 0$ the conditional variance will rise in response to an unexpected negative return. The response to a positive shock is, however, more complicated, and depends on the relative magnitudes of θ and γ. In particular, if $|\gamma| > |\theta|$ (which we find for all G7 economies), then although a positive shock will increase the

[3] Expression (3) is an ad hoc functional form designed by Nelson to capture the salient features of the data.

conditional variance, the rise will be less than that following a negative shock. The innovations z_t are assumed to follow a conditional General Error Distribution (GED). The GED allows for fatter tails than the normal distribution, which is a salient characteristic of stock market data.

We estimate the risk premium for the UK market as well as other major stock markets – US, German and French – to assess the extent to which premia are correlated across markets. We use daily returns data on the Datastream Total Markets Index[4] from 1st January 1981 to 8th December 1997 for the four countries. Three-month Euromarket rates were used as "risk-free" interest rates.

Chart 1 shows estimated UK and US risk premia. In both cases, although the risk premium was lower in 1995-96 than in previous years, it has risen again in 1997.

Chart 1

UK and US equity risk premia

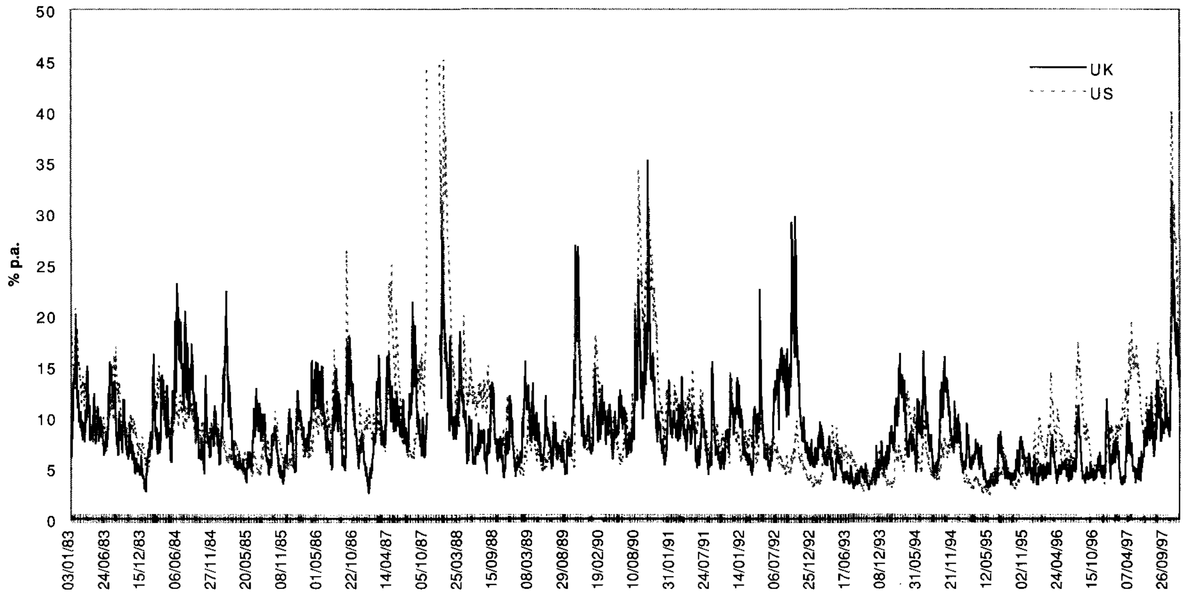


Chart 2 shows risk premia for Germany and France. They appear highly correlated, as one would expect for economies whose real and financial sectors are closely integrated. For both countries, there is no clear downward trend. But the market has been more tranquil in the 1990s, so that the risk premium is around the lower end of its range over the period.

The suspicion that the equity risk premium has fallen world-wide in recent years is not borne out by this analysis. This may be due to the high degree of persistence in volatility expectations, which means that expected volatility would fall only if actual volatility were very low for a protracted period of time.[5]

A major caveat to the above conclusion is that the standard EGARCH-M model used here assumes that the risk aversion coefficient γ is constant over time. This coefficient measures investors' willingness to bear risk. If investors have become more tolerant of risk in recent years, we would see a fall in the value of the risk aversion coefficient. For a given level of expected volatility,

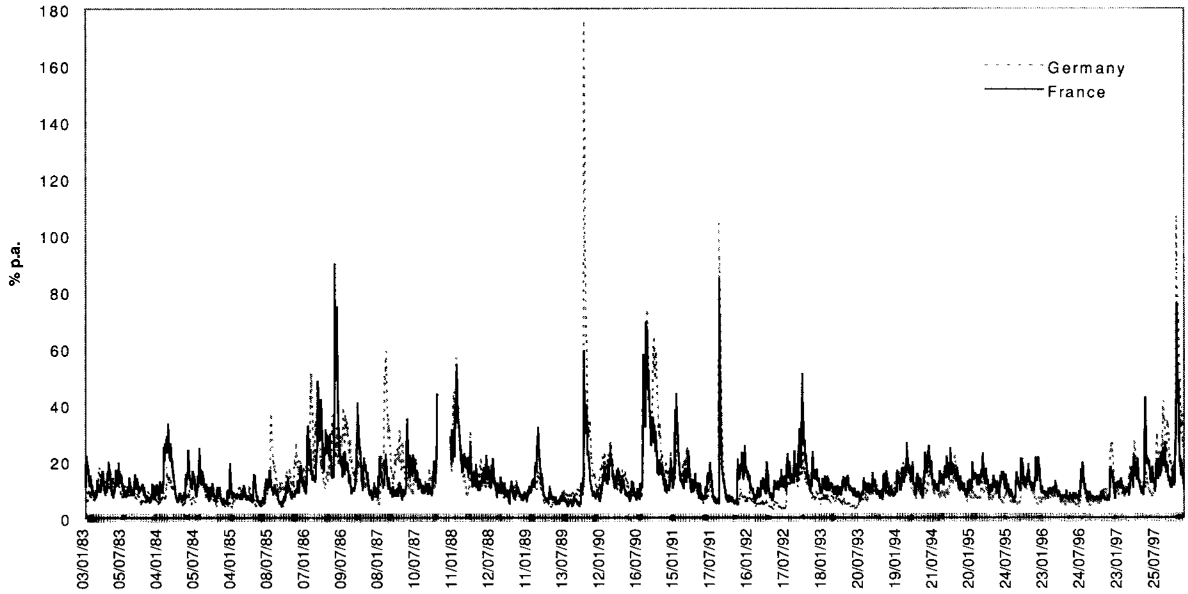
4 The Datastream Total Market Indices are broad equity indices, comprising the 1,000 largest stocks in terms of their market capitalisation for each country.

5 For example, Chen (1988) finds that the conditional variance of US stock returns can be usefully characterised as an integrated GARCH process.

this would lead to a decline in the equity risk premium. The conclusion that the risk premium has not declined may therefore be an artefact of this constancy assumption. Notwithstanding the above caveat, it appears unlikely that the current behaviour of the risk premium can explain much of the recent rise in the equity markets.

Chart 2

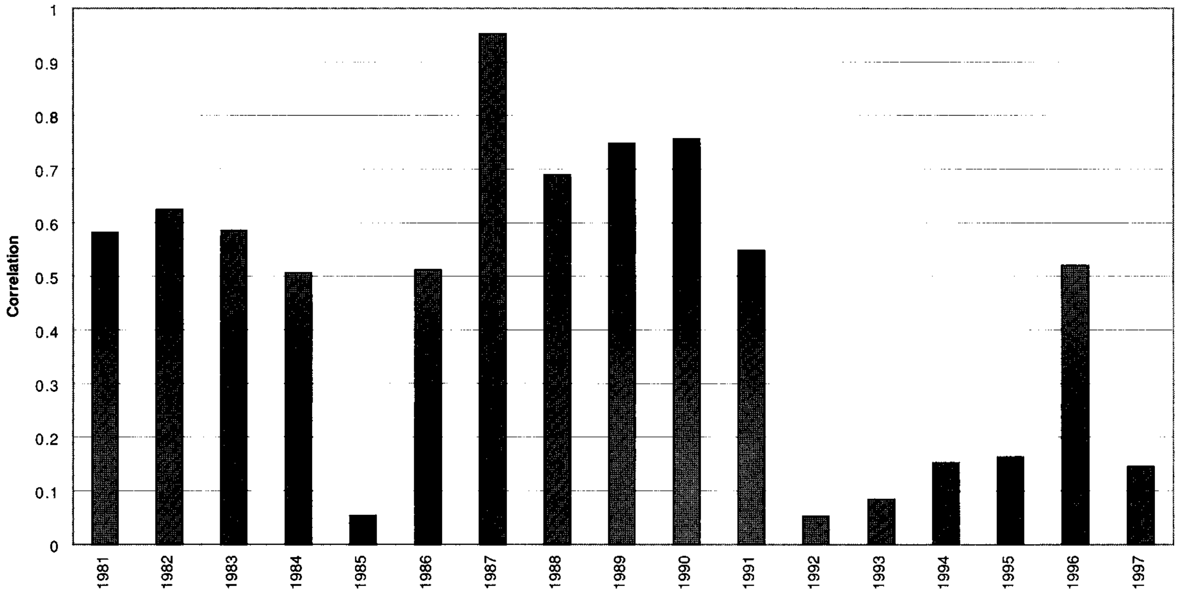
French and German equity risk premia



Looking at the correlations between equity risk premia, the UK market is very highly correlated with the United States and Canada, with full-sample correlations of around 90%. Within the G7, the United Kingdom is least correlated with the Italian market, with a correlation coefficient of 29%. However, these correlations are quite variable over time. Chart 3 shows the correlations between UK and US risk premia for each year from 1981 to 1997. The highest correlation was 95% in 1987 (when all G7 equity markets were highly correlated), but only 5% in 1985 and 1992.

Chart 3

Correlation between UK and US equity risk premia



2. Dividend growth forecasts using the Gordon Growth Model

This section looks at how we can extract market forecasts of future dividends from current equity prices. These can be compared with past dividend performance to gauge whether investors currently hold optimistic views compared with the historical performance of dividends.

The Gordon Growth Model provides a simple equation for linking the stock price to expected future dividends.

The Gordon Growth Model

The present-value formula for the price of a stock states that the current stock price P_t is the discounted present value of expected future dividends, D_t, where each dividend is discounted by the required return (or opportunity cost of capital), K, which we assume to be constant. So if we are currently in time 0, the stock price P_0 is:

$$P_0 = E_0\left[\frac{D_1}{1+K} + \frac{D_2}{(1+K)^2} + \frac{D_3}{(1+K)^3} + \cdots\right] \tag{4}$$

where E_t denotes expectations formed at time t.

Equation (4) can be simplified if we assume that dividends are expected to grow at a constant rate g. In this case, we can write all future dividends as a function of the current-period's dividend, D_0. Specifically, $D_1=(1+g)D_0$, $D_2=(1+g)^2D_0$ and so on. As long as $g<K$, we obtain the Gordon Growth formula:

$$P_0 = \frac{(1+g)}{(K-g)}D_0 \tag{5}$$

The simplest expression for g from (5) is:

$$g = \rho K - (1-\rho) \tag{6}$$

where $\rho \equiv \left(1+\frac{D_t}{P_t}\right)^{-1}$

Estimating the required return

For each individual equity or equity index we obtain the dividend-price ratio, and therefore ρ, from Datastream. The required return K, on the other hand, needs to be estimated. Equation (6) indicates that the estimate of K is extremely important for the resultant growth estimate. Although ρ is less than 1, it is generally in the range 0.95-0.99. From equation (6), this implies that a one-unit rise in K is associated with a near-one-unit rise in g. The estimate of g is therefore extremely sensitive to the estimated required return K.

We use two methods to estimate K. First, the CAPM equation, which posits a linear relation between the required return on each asset and the required return on the market portfolio, can be used to derive the required return on each stock (labelled $K1$). The resultant growth estimates are labelled $g1$ in Table 1.

Table 1

Dividend growth estimates, 4th September 1997

Sector	*R*	*K*1 (%)	*g*1 (%)	*K*2 (%)	*g*2 (%)	Past growth (%)
Financial	0.970	26.19	22.45	32.59	28.67	11.31
Mineral Extraction	0.970	19.36	15.77	25.30	21.53	10.74
General Industrials	0.963	17.37	13.03	9.91	5.84	4.82
Consumer Goods	0.969	15.22	11.62	14.38	10.80	7.13
Services	0.972	17.05	13.75	14.60	11.37	9.26
Utilities	0.959	14.99	10.29	16.54	11.76	7.65
House Building	0.967	43.80	38.13	16.94	13.14	n/a (see notes)
Construction	0.966	36.78	32.14	14.93	11.04	n/a (see notes)

Notes: r is as defined in equation (4). $K1$ is the estimated cost of capital using the CAPM as the model for required returns, and $g1$ is the associated expected dividend growth rate, calculated using equation (4). $K2$ is the average ex post return over the previous five years, and $g2$ is its associated dividend growth rate. Past Growth is the average rate of ex post dividend growth over the previous five years. Lack of historical data means that Past Growth figures are not available for House Building and Construction.

The second method is to obtain a model-free estimate of the required return by simply taking the average return on an asset over the previous five years. The intuition here is that if investors are rational, the realised return should differ from the expected return only by a white noise error term. Since positive and negative errors will cancel out over a long period of time, the average realised return should equal the (constant) expected return. This procedure results in the required return estimates $K2$, which produces the growth estimates $g2$.

Dividend growth estimates

Table 1 shows the results for the Datastream Industry-based portfolios, for data at the close of the market on Wednesday 4th September. The risk-free rate is taken to be the one-month Treasury bill yield, and the expected return on the market portfolio is calculated using the Datastream Total Market Index. The average Treasury bill yield over the past five years was 6.1%, while the average excess return on the market was 11.3%. These are the figures used in the CAPM calculations.

As mentioned above, the dividend growth forecasts depend crucially on the required return estimates. Looking at the columns headed $K1$ and $K2$, there appears to be little relationship between the two: neither is consistently higher or lower than the other, and some of the differences are extremely large. For example, for House Building the CAPM estimates $K1$=43.8%, whereas the average return over the previous five years was 16.9%. Mineral extraction, on the other hand, actually returned on average 25.3%, whilst the CAPM estimate is 19.4%. Such differences will inevitably result in differences in the growth estimates.

However, it is still possible to draw consistent inferences concerning the implications of the growth estimates for the appropriateness of current equity prices. For the Financial sector $g1$ and $g2$ are estimated as 22.5% and 28.7% respectively. Although these differ by a substantial 6%, they are both more than twice the rate at which dividends have grown on average over the previous five years, 11.3%. For Mineral Extraction, Consumer Goods, Services and Utilities, the expected future dividend growth figures are generally between one-and-a-half and twice those seen in the previous five years.

For three of the sector portfolios – General Industrials, House Building and Construction – the difference between the two growth estimates is too large for reliable inference to be drawn: the estimates range from 5.8% to 13% for General Industrials; from 13.1% to 38.1% for House building; and from 11% to 32.1% for Construction. In particular, for General Industrials the $g2$ figure of 5.8% is not too far from the past growth figure of 4.8%; but the $g1$ estimate of 13% clearly indicates stronger dividend expectations than had previously been seen.

The evidence here suggests that investors expect around twice the rate of dividend growth in the future than has been seen in the recent past. If correct, the issue is why have investors so adjusted their expectations. There are, however, two important caveats to remember. First, since the Gordon Growth Model is a steady-state model, the results will be misleading to the extent that the equity market is not in steady state.[6] Second, the results are sensitive to the estimate of the required return, K. If we have consistently overestimated the equity cost of capital, then the procedure adopted here will inevitably lead to the erroneous conclusion that the market is over-valued.

3. What type of information can we extract from equities?

The Modigliani-Miller (MM) theorem, on some strong assumptions, suggests that, the financial structure of a firm is irrelevant. Financial structure has no impact upon corporates' net worth, and should not influence their real decision making (i.e., how much to produce, invest etc). In as much as the MM theorem holds, there is no theoretical explanation why financial structure should be causally linked to firms' behaviour. The implication is that the equities are of interest only if they exhibit reliable (atheoretic) leading indicator relationships. But more modern finance theories (for example, Myers and Majluf (1984)) that stress the importance of imperfections in capital markets, suggest that these imperfections may influence the real activities of firms. The "credit view" of the monetary transmission mechanism has built upon this analysis, and suggests that at a macro-economic level, changes in equity prices may have quantitatively important effects upon corporate sector behaviour. If these credit effects matter, equity price movements may have a structural, rather than merely leading indicator, relationship with corporate sector activity.

For individuals, the theoretical underpinnings are clearer: life cycle theories suggest that wealth should be an important determinant of individual's consumption decisions, and equities (along with housing) form the main component of individuals' wealth holdings.[7]

This section reviews theory and evidence that equity price changes may have a causal impact on corporate sector activity. We couch this analysis in terms of the contribution of equities to the propagation of monetary shocks – reflecting the particular focus of central banks – but the conclusions are applicable across a wider range of shocks.

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6 Steady state means that dividend growth and the discount factor applied to dividends are expected to be constant, which underlies the derivation of the Gordon Growth Model. The Bank is currently undertaking research into using the Campbell-Shiller dividend-price ratio model as an alternative framework for deriving profit expectations from equity prices. This is a dynamic version of the Gordon Growth Model that does not require the assumption that the market is in steady state.

7 Less clear, however, are the quantitative importance of (equity) wealth effects upon consumption, and whether changes in the distribution, and form of holding equity wealth over the last decade or so have strengthened or weakened this relationship. Although an interesting issue, we leave this to one side here.

The monetary transmission mechanism and corporates: the role of equities

Theory

Traditional views of the transmission mechanism, as embodied in IS-LM analysis, focus upon the power of monetary policy to change the real cost of capital in the short-run. Changing the cost of capital alters the returns from savings and investment, and so the level of real output. Such models typically do not model the equity market. Implicitly any change in equity prices are viewed as an endogenous response to the changes in real activity brought about by the cost of capital channel: in these models equity prices may change, but there is no additional effect from this over and above that brought about by the change in the bond prices.

This traditional model has been extended to include additional assets markets (e.g. Brunner and Meltzer (1972)). The extended models imply a richer transmission mechanism for monetary shocks, with the output effect now depending on the interaction of multiple (two in the case of Brunner and Meltzer) asset markets and output.

But all of these traditional views implicitly assume that capital markets are perfect. By contrast, the "credit view", which has been developed by Bernanke, Gertler, Gilchrist and others over the last ten years or so, stresses the contribution of capital market imperfections to the transmission mechanism.[8] The focus has been on the role of information gaps in capital markets, rather than the more tangible distortions brought about by the tax system and transactions costs.

Information gaps arise because it is difficult and costly to monitor the state of firms. This can create principal-agent problems between both debt and equity holders and managers. There is a substantial literature which details the precise nature of these problems (see Gertler (1988) and Gertler and Gilchrist (1993) for surveys). Its main conclusions are well known. First, it is in the interests of lenders to place restrictive covenants on firms' behaviour, to ensure that firm managers do not act against the lenders' interests. Second, the restrictiveness of these covenants is likely to increase as the debt to equity ratio of a firm increases. Third, firms will have to pay a premium for new equity issues, if the attempt to issue new equity is likely to be interpreted as signalling that management believe prevailing market value of equity is unwarrantedly high. This is likely if managers have more information about the state/value of a firm than shareholders,[9] and alternative forms of finance are available.[10]

These arguments underpin the famous "pecking order theory of finance" (Myers and Majluf (1984)) which suggests that when such information gaps are germane, firms will find that internal finance is cheaper than external finance, and within that new debt will tend to be cheaper than new equity. These theories suggests two further channels of monetary transmission over and above the simple cost of capital channel (Bernanke and Gertler (1995)).

First, there will be a *balance sheet*, or *net worth* channel. This rests upon the assumption that the size of the premium attached to external over internal finance – the external finance premium – will increase as the net worth of a firm decreases. The intuition is that a stronger financial position reduces the potential conflict of interest between a manager and the debt holder. For example, as the

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8 The seminal article which underpins much of the credit chanel literature is Stiglitz and Weiss's (1981) analysis of equilibrium credit rationing.

9 Managers will have an incentive to issue shares if the current share price over-values the firm, as equity finance will then be cheap; conversely, managers have no incentive to issue shares when the current share price under-values the firm. This is an example of the famous "lemons" problem.

10 Some firms, for instance high growth firms, may not have to pay a premium as they credibly argue that current cash flows would be unable to finance expansion and that (sufficient) debt is not available.

net worth of a firm in financial distress improves there is less incentive for managers (assuming they maximise equity-holders returns) to undertake negative net present value projects which offer a small probability of a very large return.[11]

Bernanke and Gertler (1995) argue that equities may play a causal role here. Monetary shocks are likely to change equity prices, and in so doing will change the value of borrowers' collateral, which in turn is likely to change firms' perceived credit-worthiness and so the premium charged on new loans.[12] The result is the so-called "financial accelerator" (Bernanke, Gertler and Gilchrist (1994)): the initial impact of, for example, a negative shock to production and investment from the cost of capital channel will reduce net worth. This will increase the external finance premium, amplifying the cost of capital effect upon activity.

The importance of this accelerator is likely to be greater in recessions than in booms. The reason is that there is a lower bound of zero on the external finance premium. When the economy is in good shape, borrowers may charge a lower premium than during recessionary periods, on the basis that the general riskiness of lending has fallen.

The second additional channel of transmission is the *bank lending* channel. This will be relevant if there are agents for which the only form of external finance available is bank debt. This may occur if the information gap is especially significant for some classes of agents (for example small firms and individuals) and banks have specialised at gathering and processing information about such borrowers. If this is the case, and if monetary shocks alter the relative cost of loanable funds for banks, then the transmission of the monetary shock through to bank-dependent borrowers will reflect both the basic cost of capital element, and the change in the relative price of bank debt.[13]

Evidence

The importance of the balance sheet effect is likely to vary across classes of firm. The creditworthiness of mature firms with stable earnings is easier to assess for instance, than that of high growth, new technology firms (for example, IT software houses). Accordingly, empirical studies have tended to investigate whether there is evidence of cross-sectional or time series variation across classes of firms' behaviour that is consistent with a balance-sheet channel. Schianterelli (1996) provides an excellent summary of the literature.

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11 Imagine a firm has very low net worth and is likely to go bankrupt. If bankruptcy occurs the equity holders will receive nothing – debt holders get the first claim on bankrupt companies. Thus, there is an incentive to undertake very risky projects, on the off chance that they will generate sufficiently large returns to make the company profitable again. By contrast, if the net present value of such risky projects is actually less than zero, then the debtholders' expected return will be reduced further by investment in risky projects. And conversely, if there is a very safe project, which is likely to make money, but leaves the company still insolvent, there would be no benefit to equity holders from investing in the project, even though it would increase the expected return to debt holders.

12 Suppose a monetary shock (rise in interest rates) decreases equity prices. If a firm's debt is fixed rate then its value will be unchanged and the debt/equity ratio will rise. Alternatively the debt may be floating rate; in this case the value of debt interest payments and the debt will rise. Now the rise in the debt/equity ratio will be even greater, as equity value will have fallen and debt value will have risen. Thus whatever form existing debt takes, potential new lenders will observe a fall in the available collateral for new loans.

13 This second element of the credit channel is more controversial. For example, Romer and Romer (1990) argue that changes in the US regulatory structure during the 1980s increased the liquidity in financial markets and made it easier for US banks to raise wholesale funds, making the supply of loanable funds to banks more elastic. In the United Kingdom, the share of wholesale deposits in M4 has increased from an average of around 20% between 1983-85 – the pre "Big Bang" period – to around 30%, over the last three years; consistent with the notion that wholesale funding has become easier for banks.

Firm level studies have tended either to estimate cross-sectional investment equations directly, and test whether financial factors have a role in explaining the investment behaviour of constrained firms, or to estimate first order conditions for investment – Euler equations – and test whether these are violated for constrained firms. A common strategy amongst the papers that directly model investment has been to assess whether specification failures in Tobin's q can be explained by financial variables. The rationale comes from Hayashi (1982), who showed that Tobin's q will provide the optimal investment rule for firms only if capital markets are perfect, and if there are (known) installation costs associated with investment.

One of the first studies was by Fazzari, Hubbard and Peterson (FHP) in 1988. They found that financial structure variables play an important role in explaining investment behaviour across different classes of publicly quoted manufacturing firms in the United States. They divided their sample of firms into three categories, according to their dividend payout ratios. Those firms who pay the least dividends are likely to have exhausted available internal funds, and so have to rely on external funds to finance investment (Myers and Majluf (1984)). By contrast, firms which pay high dividends are likely to have sufficient internal funds to finance investment, or do not have to pay a significant external finance premium – perhaps because they are well know firms operating in mature industries. If capital market imperfections are unimportant then, as discussed above, variations in Tobin's q should be able to account for firms' investment. Consistent with this notion, FHP found that the investment behaviour of the high dividend payers could be adequately explained by Tobin's q ratio, but that financial factors (cash flow) were an important additional determinant of investment for the lowest dividend payout class of firms.[14] Many subsequent studies using different proxies for financial factors have reached similar conclusions (Gertler and Hubbard (1988), Whited (1992), and Hoshi, Kashyap and Scharfstein (1991)).

Studies of US firms that adopt the Euler equation approach generally reach similar conclusions. For example, Hubbard and Whited (1995) find that the over-identifying restrictions for the Euler equation are rejected for financially constrained firms (those that pay low dividends) but are not rejected for financially unconstrained firms. Cross-sectional evidence in the United Kingdom is slightly less compelling. For example, Bond and Meghir (1994) estimate Euler equations for constrained and unconstrained firms. While they find that financial factors are unimportant for unconstrained firms – consistent with the theory – they find that the violation of the Euler equation is wrongly signed for constrained firms, in the sense that increases in cash flow are correlated with falls in investment.

More recent work at the Bank of England (Small (1997)) analyses whether cash flow has a significant positive effect upon inventory investment. The study analyses the investment behaviour of 605 UK-quoted firms over 1977-94 whose prime business activity was manufacturing. Inventory investment is modelled as a function of the lagged stock of inventories, current and lagged sales and a cashflow term.[15] The importance of cash flow is then investigated, with the firms divided into constrained and unconstrained groups according to four characteristics: dividend behaviour, interest cover, firm size and the current ratio.[16]

The study finds that firms' current cash flow has a significant positive effect upon inventory investment. However cash flow appears to matter for both constrained and *un*constrained

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14 One objection to this type of study is that measures of Tobin's q - average q - may actually be a very noisy measure of the true shadow value of marginal capital expenditure, so that tests for incremental explanatory power from financial variables is weak. As Hayashi (1982) also demonstrated average q will only always equal the economically important marginal q when firms are price takers, and have technology which exhibits constant returns to scale.

15 As in common in panel data analysis, firm and specific effects are also allowed for.

16 Interest cover is defined as the ratio of interest payments to operating profits and the current ratio as the ratio of current assets to current liabilities.

firms (under each criteria). This finding is puzzling as it suggests financial structure matters even when there are no financial constraints. Some more limited support for the credit view is provided, however, in that the size of the cash flow effect upon investment seems to be greater for constrained firms than unconstrained firms, and the difference appears statistically significant.

An alternative – but complementary – time series approach was adopted by Gertler and Gilchrist (1994). They investigated the sensitivity of output and inventory investment by manufacturing firms to monetary shocks through time. They split their sample into "small" and "large" firm sub-samples, and found a greater sensitivity in the behaviour of small firms, even after controlling for the variation in firms' sales. This provides indirect evidence of an external finance premium leading to differential monetary policy effects across differing classes of firm.

Similar analysis has been carried out at the Bank (Ganley and Salmon (1997)). This work has focused on the disaggregated effects of monetary policy shocks on the output of 24 sectors of the UK economy. The principal aim of the analysis was to provide stylised facts about the sectoral responses to unexpected changes in monetary policy. However, it also provided indirect evidence about the underlying nature of the transmission mechanism by suggesting that the effects of unanticipated monetary policy tightenings are unevenly distributed across sectors of the UK economy. As might be expected, sectors such as construction show a sizeable and rapid decline in output, whereas others, like services, show a much more muted reaction. Manufacturing as a whole also responds quite sharply to a monetary tightening, but some large industrial sectors, notably utilities, show a subdued reaction. Moreover, the 14 sub-sectors that comprise manufacturing also exhibit diverse responses to a monetary shock. The paper shows that the pattern of these sectoral manufacturing responses seems correlated with the size characteristics of the firms in each sector.[17] In particular, sectors which mainly comprise "small" firms tend to exhibit a stronger reaction to monetary shocks than sectors that mainly comprise "larger" firms. This result is consistent with a "credit view" of the transmission mechanism, in as much as the small manufacturing firms experience greater variation in their external finance premium. But of course, other factors could lie behind this pattern.

4. Extracting information from equities

The evidence presented in the last section suggests that equities may have a structural, as well as leading indicator, relationship with firms' investment. Further, it suggests that the importance of financial factors upon firms' activity is clearly going to vary across types of firm, and the state of the business cycle.

The difficulty is that concluding that equities may have structural importance is not akin to identifying a structural model that can be estimated to test this hypothesis. The credit view of the transmission mechanism in particular does not offer a unified alternative to traditional views of the transmission mechanism. Rather, it just suggests ways in which the traditional view may be deficient.

From a modelling perspective this points to an atheoretic approach. This can, at the very least, help answer the most basic question as to whether equities contain leading information for the rest of the economy, regardless of whether this derives from structural relationships.

This section presents the results from some preliminary VAR work that is in the spirit of this approach, and then discusses how it might be extended.

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[17] Because of the historic importance of manufacturing in the United Kingdom, more detailed data are available for this sector than for the rest of the economy, even though its aggregate importance has declined. Hence, it was possible to carry out the "size characteristics" analysis for only the manufacturing sector.

Following work in the United States by Lee (1992) we have estimated a small (non-structural) VAR model with four variables: real equity returns, real interest rates, growth in industrial production, and inflation. The "causal" ordering used in the VAR model is as follows: real equity returns, real interest rates, growth in industrial production and inflation. The only deterministic component in the VAR model was a constant, and we used six lags of the variables on monthly data from April 1988 to December 1994. The real interest rate was the return on the Treasury bills less the inflation rate, computed as the monthly change in the Retail Price Index. Real equity returns were computed as the monthly return on the value-weighted FT-All Share Index less the inflation rate. Real activity in the economy was proxied by the growth in industrial production.

The results from this simple VAR model are provided in Table 2. The main points to note are:

1) Real equity returns do not appear to be exogenous, in the sense that after 24 months over 30% of their variation is explained by the three other variables in the system. Of the other variables the real interest rate is the most important, consistent with the notion that monetary shocks have a significant influence on equity prices.

2) Equities do appear to have some incremental information in terms of forecasting real activity. After 24 months equities account for 8.6% of the forecast error variance in real activity, compared with 12.4% for real interest rates.

3) After 24 months, only about 7% of the variation in inflation can be attributed to real equity returns, while innovations in real interest rates explain almost 30% of the variation in inflation.

3) After 24 months, almost 67% of the variation in real interest rates is explained by innovations in inflation.[18] After a similar period, real equity returns only account for 14% of the variation in real interest rates.

Table 2

Simple VAR model results (in percentages)

	By innovations in:			
	Real equity returns	**Real interest rate**	**Industrial production growth**	**Inflation**
Variable explained (after 24 months):				
Real equity returns	69.9	18.6	4.6	6.9
Real interest rate	14.4	8.6	10.6	66.1
Industrial production growth	8.6	12.4	69.8	9.3
Inflation	6.6	29.4	12.2	51.8

Note: Due to rounding errors, the rows may not add to 100%.

Lee (1992) obtains similar results. In particular, US real equity returns also appear to have some Granger causal information for real activity, as measured by growth in industrial production. Roughly 11% of the variance in real activity can be ascribed to real equity returns. Impulse response analysis shows that the response of real activity to shocks in real returns is strong

[18] A complication of interpreting this result arises because we have used a backward-looking measure of real interest rates by using inflation outturns.

and positive for the first 12 months, after which it tapers off. US equities appear to have even less information for inflation than UK equities.

The results for the United Kingdom are very preliminary; no attempt has been made to gauge the robustness of the reported results to the choice of lag length or alternative causal ordering in the VAR model. In addition, the analysis does not use a particularly long span of data, and considers only one data frequency, monthly. But, more fundamentally, we intend to extend the form of the VAR in a number of ways.

(i) Disaggregation

Disaggregated analysis has clear potential to isolate clearer, and distinct links between equities and activity. We propose to construct equity returns for industry sectors (either equally weighted or value weighted) which would reflect the general trend of share price movements in the specific sector. This could then be combined with more disaggregated macroeconomic data, in order to construct a range of sectoral VARs. Another aspect of the disaggregated analysis would be to consider the contribution of various sectors to the aggregate growth outlook implied by the equity market. This would provide an indication of the degree to which growth is likely to be balanced or concentrated in, say, exports, as in the early stages of the most recent recovery.

(ii) Sub-period analysis

One implication of Section 1 is that the external finance premium, and possibly the importance of equity price changes, will vary with the state of the cycle and monetary policy. Sub-period analysis (as in Titman and Warga (1989)) would help establish whether this is the case. In terms of monetary policy, estimation over distinct sub-periods (for example, post ERM) might be more appropriate.

Separately, there is the question of the forecast horizon, i.e. how many periods ahead would we expect current equity returns to forecast real activity? In theory, movements in equity returns reflect agents' expectations over an infinite time horizon. But, in practice, the time horizon will be shorter as a result of the effects of discounting and, possibly, the short-termist nature of the equity market.

(iii) Expectations extraction

A complementary approach to examining the information content of equity prices would be to extract expectations of future dividends and discount rates contained in equity prices, and then consider how these relate to future real activity. This approach may be preferable for a number of reasons. First, it is likely that, in practice, equity returns will turn out to be a noisy indicator of future macroeconomic activity. By considering the expectation variables directly, it may be possible to remove (or reduce) noise, and hence avoid this problem. Second, this approach would also have the advantage that it would be possible to ascertain the relative importance of expectations of dividend growth and discount rate changes for predicting real activity.

For our purposes, the first step would be to use the VAR model employed by Campbell and Shiller (1989), Campbell and Ammer (1993) and Paisley (1995) to generate expectations of future dividends and discount rates. The second step would then involve using these expectations variables, instead of (raw) equity returns data, in our original VAR model to consider the leading indicator properties of these variables. Note, however, that these expectations variables would still contain some noise.

5. Information from options

In this section we focus on traded options on the FTSE 100 Index to obtain ex-ante information about future market moves, which in turn might have implications for consumption and investment behaviour.

Option markets provide a richer source of information than is available from the futures market. Unlike futures markets, which only provide information about the expected future *level* of the market, implied PDFs derived from the options market tell us what probability agents attach to all possible values of the index at some terminal date. Consequently, by focusing on the information embedded in implied PDFs, monetary authorities can reach a more comprehensive assessment of overall market sentiment.

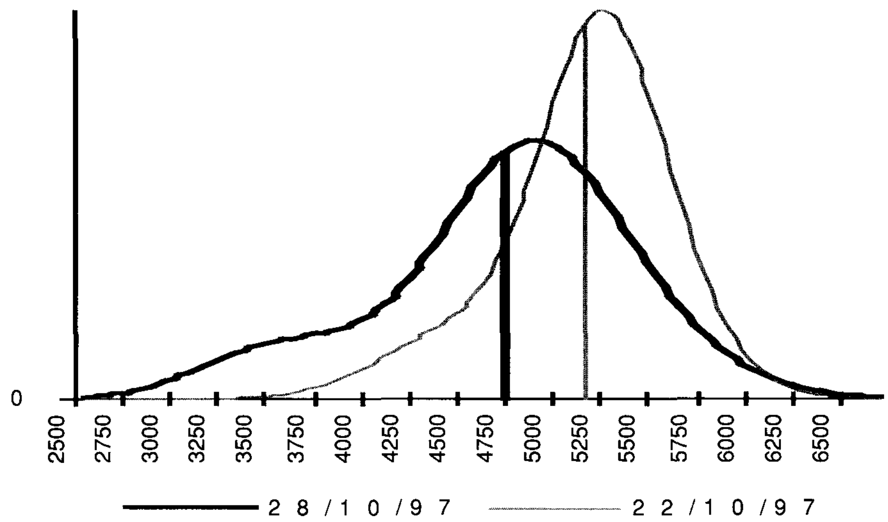
However, before looking at changes in the PDFs around specific events, it may be instructive to outline the method used in the Bank to extract them. Briefly, an option price is assumed to be equal to the present value of the discounted probability-weighted future payoffs to the option. What we are interested in is finding out the probabilities attached to each possible payoff at the maturity of the option. We use a non-linear optimisation routine (Powell) to minimise the squared difference between observed and theoretical option prices to recover the set of probabilities consistent with the observed option prices. To obtain these probabilities we need to assume a particular distribution for the underlying instrument at option maturity. It is well known that asset price distributions are not log normal, but that they are close to log normal. A distribution that is close to, but not exactly, a log normal distribution can be closely approximated by a linear combination of two log normal distributions – a mixture distribution. Our optimisation technique recovers the mean and the variance of the two log normal distributions, and the relative weights attached to the two distributions. The overall shape of the mixture implied distribution is entirely determined by 5 parameters. Bahra (1997) contains a detailed technical description of the technique used in the Bank.

It is worth bearing in mind that the PDFs extracted from option prices are risk-neutral, rather than the market's subjective distribution of expectations. This is because options are priced using a risk-neutral distribution. Rubinstein (1994) shows how one can link the risk-neutral and the subjective distributions for a "representative" risk-averse investor. Some preliminary work along the lines suggested by Rubinstein has been conducted in the Bank and the upshot of the research is encouraging: the risk-neutral and subjective FTSE 100 implied distributions are qualitatively similar.

Chart 4

FTSE 100 PDF, December 1997 contract

Probability density (% probability per 10 b.p.)



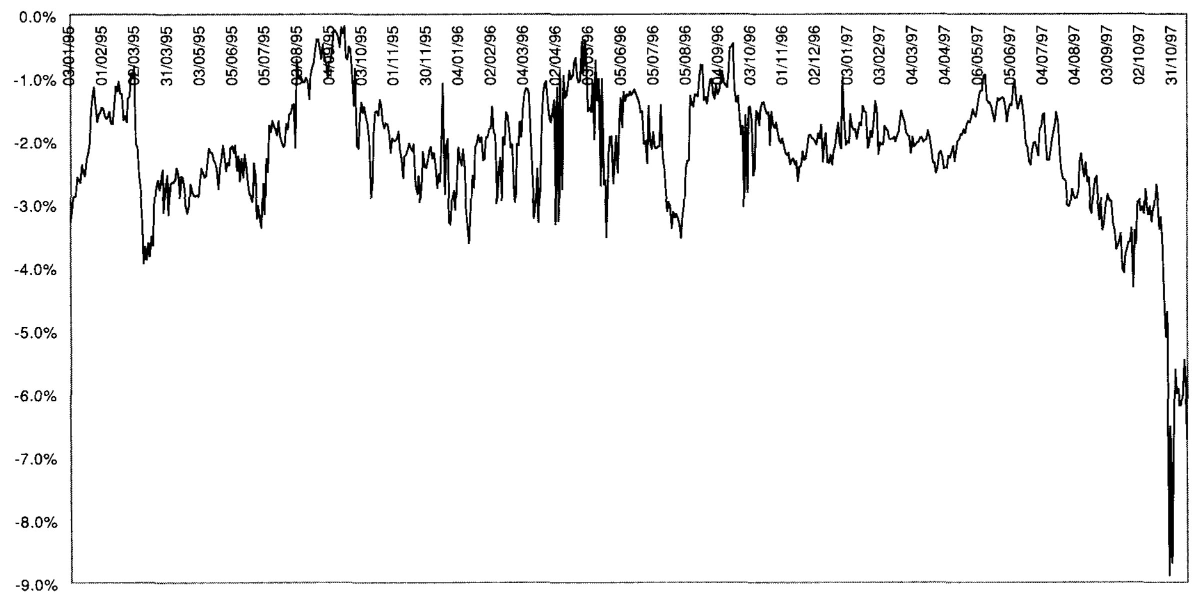
To illustrate the usefulness of this technique in providing information about market sentiment which is not adequately revealed in the level of the index, we focus on a specific event in October of this year. Chart 4 shows the implied risk-neutral FTSE 100 PDFs on 22nd and 28th October for the December contract. The turbulence in the Hong Kong market spilled over into other world equity markets on 23rd October. It is immediately apparent that there is more probability mass in the left-hand tail of the distributions, implying negative skewness. Intuitively, this means that the market attaches a higher probability to further large falls relative to large rises. This negative skew is a common feature of the UK equity options market and has existed since 1990.[19] Moreover, it is clear from Chart 4 that the negative skewness became much more pronounced after the sharp fall in the FTSE 100 between the two dates. This large increase in negative skew was also apparent in the March and June 1998 contracts.

What implication does the negatively skewed implied FTSE 100 distribution have for monetary policy? One possibility is that monetary authorities should be less concerned about the possible wealth effects on consumption of a rise in the stock market. This is because, to the extent that the strong negative skew in the implied distribution reflects the fact that agents are attaching a large probability to a significant market correction, the impact of wealth effects on consumption is likely to be somewhat subdued.

To gain a longer term perspective on how agents' expectations about the future level of the FTSE index has changed, we look at the time series of skewness of the implied PDFs. At this point it is worth pointing out that the time series of moments that we recover from the implied PDFs cannot be compared from one day to the next (or over longer measurement intervals). This is because option contracts have a fixed time to maturity, and so, as the option contract nears maturity, agents' uncertainty about the price of the underlying asset on which the option is written tends to decline. In other words, the time-series of implied moments exhibit a time trend. The presence of a time trend in the implied moment makes it difficult to ascertain the extent to which day-to-day changes in the implied statistic are due to news, or simply a consequence of the fact that the option is closer to maturity. In the analysis that follows, we strip out the effect of declining time to maturity from the implied statistic.

Chart 5

Skew in FTSE 100 PDF



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19 The US market is also characterised by negative skewness and, if anything, the negative skewness in the US market is even more pronounced and has existed since the 1987 crash.

Chart 5 shows how adjusted skewness in the FTSE 100 PDF has evolved since 1995. Although the skew has been negative for the whole period, the turbulence in October caused a dramatic increase in negative skew from around -3% to around -7%.

The message from the PDFs is that the market has been consistently pricing in the possibility of a correction. There is some indication that the market thinks that the likelihood of such a correction has increased over the year, particularly since the turbulence in Asian and Latin American equity markets. It seems plausible that the impact of wealth effects on consumption may be lower than if the market were at the same level, but with a more symmetric distribution.

Summary

In this paper we discuss techniques that enable us to extract information about the equity risk premium and dividend growth expectations from the market. From a policy perspective, the equity risk premium is of interest because it is an important component of the cost of capital which, in turn, is an important determinant of investment expenditure in the economy. Besides, changes in the risk premium have implications for the level of the market. Dividend growth expectations are closely correlated with profit expectations; therefore, by focusing on the former, monetary authorities may be better able to assess inflationary pressures in the economy.

We have looked at the role of equities in the transmission mechanism. Both the traditional and "credit" views of the transmission mechanism are discussed. In particular, we review the theory and empirical evidence which suggests that equity price movements may have a causal impact on the corporate sector.

With regards to extracting information from equity markets, some preliminary work with a VAR model suggests that equities may contain information about the real economy. It is likely that the components of returns, such as expectations of future dividends and discount rates, may prove more useful leading indicators, but much work remains to be done before we can be more certain.

Option markets enable us to extract the probability agents attach to all possible levels of the market at some terminal date and so allow a more comprehensive assessment of market sentiment. More work needs to be done to see whether the moments of the implied distribution, such as skewness, can predict future market moves and whether they are significant determinants of consumption and investment behaviour. At the Bank we are currently investigating these issues.

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Information content and wealth effects of asset prices – the Austrian case

Heinz Glück and Richard Mader*

In the framework of Austria's monetary and exchange rate policy, asset prices traditionally did not play a prominent role. Over long periods, asset formation, to a very large extent, took place only in the form of savings deposits via the banking system. Capital markets did not develop sufficiently to be a factor of great concern or an interesting source of information. Consequently, also wealth effects in the transmission process of monetary policy could be largely disregarded. Recent years, however, brought increasing discussion and research on this topic at the international level. Thus, in this paper we try to approach two questions which seem to be of major importance in this context, namely the information content of financial asset prices and the weight of wealth effects in the transmission process.

1. The information content of the term structure

1.1 Introduction

In recent years empirical research has increasingly focused on the interdependence of asset returns, inflation and real activity. In this respect, yield curve spreads and stock prices in particular are thought to contain valuable information and, therefore, are usually taken into account in economic forecasting.

In principle, policymakers can make use of these financial indicators in implementing monetary policy. Yield curve spreads can be used as a benchmark for macroeconomic forecasts, for example. On the one hand, if the indicators coincide with the model forecasts, confidence in the model results is enhanced. On the other hand, in the case of inconsistencies between financial indicators and model results, a review of the assumptions or the structure of the model may be necessary. Additional advantages are that yield curve data are available in real time and are not subject to revision.

For the United States, strong evidence exists that the steepness of the yield curve is a good predictor for real activity[1] (Estrella and Hardouvelis (1991), and Estrella and Mishkin (1996)). In other industrial countries, in particular the EU economies, the evidence is mixed. Good forecasting properties of the slope of the yield curve have been identified for Canada and Germany (Plosser and Rouwenhorst (1994), Estrella and Mishkin (1995), and Bernard and Gerlach (1996)) and to a lesser extent for the United Kingdom (Estrella and Mishkin (1995), and Bernard and Gerlach (1996)), Italy (Estrella and Mishkin (1995)) and France (Davis and Fagan (1995), and Bernard and Gerlach (1996)).[2]

Moreover, Estrella and Mishkin (1996) concluded that the information content of yield

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* The views expressed are those of the authors and not necessarily those of the Oesterreichische Nationalbank (OeNB). The authors are grateful to Helmut Pech and Peter Mooslechner for helpful comments.

1 In the United States the best predictability was found for 6 and 8 quarters ahead (Estrella and Mishkin (1996)).

2 Estrella and Mishkin (1995) did not find any information content of the slope of the yield curve for predicting recessions in France.

curve spreads is far greater than that of stock prices. However, predictions for real output in the United States were found to become more accurate, if the stock price index was included as a regressor. Thus, stock prices seem to contain information which is not reflected in the yield curve spread and are, therefore, thought to be useful in forecasting recessions and/or recoveries.

Empirical work has also concentrated on analysing the predictive power of yield curve spreads in forecasting future inflation changes. The results indicate that yield curve spreads contain information about future inflation in the United States,[3] although the predictive power is found to be weaker than for real activity (Fama (1990), Jorion and Mishkin (1991), and Estrella and Mishkin (1995)). For the EU countries, empirical results vary but are generally less significant than for the United States (Mishkin (1991)). Davis and Fagan (1995), for example, showed that out of the four largest EU countries, spread variables performed best in the United Kingdom and weakest in France. Empirical analysis of Jorion and Mishkin (1991) points to a limited information content of spreads in Germany. However, Gerlach (1995), using a longer sample period and different spreads, found that German spreads do contain considerable information about future changes in inflation.

Researchers originally used simple or multiple regressions to analyse the information content of the term structure for economic activity and inflation (Estrella and Hardouvelis (1991), Plosser and Rouwenhorst (1994), and Fama (1990)). More recently Vector Autoregressive (VAR) models have been employed to shed light on the forecasting abilities of the term structure for inflation and real activity (i.e. Davis and Fagan (1995), and Canova and De Nicolo (1997)). In principle a VAR – if appropriately specified – approximates the data generating process of a vector of variables and allows to take into account the interdependences between the variables.

1.2 Importance of financial indicators in the Austrian monetary policy framework

In Austria, financial asset prices traditionally have not played a prominent role as monetary indicators. This partly reflects the fact that Austrian capital markets – against the background of the high degree of monetary wealth formation via the universal banking system[4] – have not developed sufficiently in order for asset prices to carry a high information content. Thus, in the implementation of monetary policy, which for more than 17 years has been oriented towards holding the schilling stable vis-à-vis the Deutsche mark, financial indicators, such as yield spreads, have not been given great importance.

In addition, the lack of data – especially the absence of long time series – makes quantitative analysis difficult, which is a major reason for the absence of empirical research of these issues.

However, in spite of these problems we will attempt to analyse whether financial asset prices can provide information and be used as indicators in monetary policy decision making in Austria. In this context, we will concentrate on the information content of yield curve spreads in forecasting economic activity and future inflation.

1.3 Data and methodology

Initially, a regression technique was applied (see Appendix 2) to predict future economic activity, according to the specifications of Estrella and Hardouvelis (1991) and Plosser and

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3 In the United States the forecasting power seems to be best for more than 6 months ahead (Mishkin (1990)).

4 The financial structure plays an important role in this context. The Austrian financial system is dominated by universal banks, which provide the major part of external financing for enterprises. Small and medium-sized companies, which dominate the enterprise sector, use credit financing as their principal source of external finance. At the same time, a substantial amount of monetary wealth is held with banks (i.e. in savings accounts).

Rouwenhorst (1994). By using the cumulative growth rate as the dependent variable, long-term economic relationships can be analysed. Thereby, the cumulative industrial growth rate was regressed on the term spread (SM-VIB)[5] on the one hand and on the term spread and a short-term interest rate (VIB)[6] on the other. A Newey-West procedure was used to correct the standard errors.[7] The analysis was based on monthly data between 1983:1 and 1995:12. When the term spread was used as the only regressor, coefficients were significant for all time horizons, with the exception of the 6-month span. When the short-term interest rate is added, results were significant for 18, 24 and 36-month horizons. However, further analysis shows that the Newey-West correction procedure is not sufficient so that the results appear to depend considerably on the method chosen. The inclusion of an AR(3) model for the error terms in the equation reveals almost no significant results (apart from the 12 and 18-month horizons for the regression suggested by Plosser and Rouwenhorst). Regressing changes in inflation on the term spread (see Fama (1990)) produces similar results. When an AR(1) model was included in the equation all coefficients became insignificant.[8]

The following analysis is based on a Vector Error Correction Model (VEC). This methodology facilitates an analysis of the interdependences between the included variables and – above and beyond pure VAR models – to take into account possible relationships among the levels of the variables.

The model comprises the following two variables:

- Real activity, which is measured by the industrial production index (IP)
- Inflation, which is measured by changes in the consumer price index (INF).

The financial variables used in the analysis are the slope of the Austrian yield curve as defined above; the slope of the German yield curve as measured by the difference between the yield on long-term domestic government bonds in the secondary market and a short-term money market interest rate (BMB-B3M); and the foreign yield spreads as measured by the difference between the yields on domestic and foreign 10-year government bonds. (Besides the Austrian/German yield differential (SM-BMB) the German/US yield differential (BMB-BMU) is included.)

In selecting the variables, the following considerations were emphasised. The choice of the yield spreads was based on their presumed macroeconomic relevance. The German yield spread and the Austrian/German yield differential were chosen because of the strong economic links between the Austrian and German economies (reflected also in the exchange rate target of the OeNB). The yield differential to the US is assumed to cover any potential relationship with the German (and therefore indirectly with the Austrian) rate. As far as the domestic yield curve is concerned, the lack of data confined the analysis to the spread between the 10-year government bond yield and a 3-month money market rate. Thus it was not possible to analyse the usefulness of different spreads for predicting inflation and economic activity. Gerlach (1995), for example, found that in Germany the medium-term range of the yield curve and, in particular, spreads vis-à-vis 2-year rates are most indicative of future inflation.

The details of the data sources and time series employed are given in Appendix 1. The sample covers monthly data from 1983:1 to 1995:12. The analysis is based on monthly – instead of

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5 The slope of the Austrian yield curve as measured by the difference between the yield on long-term domestic government bonds in the secondary market and a short-term money market rate.

6 Interest rate on 3-month interbank loans.

7 Estrella and Hardouvelis and Plosser and Rouwenhorst applied the Newey-West procedure to correct the standard errors, since the dependent variable is a long-horizon growth rate and, therefore, the error terms in the regressions exibit substantial serial correlation.

8 When a Newey-West correction is used, all results are significant (apart from the 36-month horizon).

quarterly[9] – data, so as to maximise the number of observations. A sufficiently large sample seems, above all, important to insure reliable results, although there may be the potential problem of noise in the data.

In the VEC all changes of variables are regressed on changes of the variable itself and on lagged changes of all other variables contained in the system. In addition, the VEC includes a cointegration term that conveys information about the levels of all variables in a special restricted form.

Unit root tests were conducted and unit roots were found in all the variables. Thus, the precondition of a VEC model is fulfilled with all time series being integrated of order 1. In calculating the VEC no deterministic trend in the data was assumed. According to the Akaike and Schwartz criteria, two lags were found to be the optimal lag structure.

A Johansen Test (see Appendix 4), which can be interpreted as a multivariate unit root test, was performed to check for the existence of stationary linear combinations of variables. The test found one cointegrating relation at a 5% significance level.

1.4 Empirical results

The results of the VEC analysis (see Appendix 3) indicate that, apart from the dependence of inflation on industrial output, all other variables are autonomously driven. Thus, the analysis does not point to any predictive power of the considered domestic or foreign yield spreads for output or inflation. The changes in inflation significantly depend on the levels of all other variables, as they are closely tied by the error correction term.

This finding contrasts with the empirical results generally found for the United States, Canada and for – though on a minor scale – some of the larger EU economies, in particular Germany (Davis and Fagan (1995), Bernard and Gerlach (1996), and Canova and De Nicolo (1997)). In several smaller economies (more comparable to Austria) predictive power was attributed to yield curve spreads, but no in-depth research has been conducted for such countries.[10]

The inclusion of stock prices does not change these results. Output and inflation are not significantly affected by the stock market. This confirms the finding of Canova and De Nicolo (1997) for some larger European economies (Germany and the United Kingdom). In the United States, by contrast, stock prices were found to add information not contained in yield curve spreads (Estrella and Mishkin (1996)).

1.5 Conclusions

The limited relevance of yield curve spreads for predicting real activity and inflation in Austria – in particular compared to the larger economies such as the United States, Canada or Germany – may be traced to the following factors.

First of all, the lack of data, which precludes a more comprehensive and in-depth analysis of the issues discussed (such as analysing different yield spreads or longer lag structures).

Second, partly as a result of the small size of the market, interest rates did not respond strongly to market forces before the late 1980s, when liberalisation and deregulation of Austrian financial markets started to accelerate. The size of the bond market was limited, and the market was

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9 However, analysis with quaterly data did not substantially change the results.

10 Yield curve spreads were found to be significant for predicting real activity in Belgium and the Netherlands (Davis and Fagan (1995), and Bernard and Gerlach (1996)) and for forecasting inflation in Belgium (Davis and Fagan (1995)).

not fully developed over much of the sample period. Increasing integration of financial markets might improve the forecasting power of yield spreads.

Third, the financial market structure may partly explain the poor forecasting performance of yield curve spreads. Institutional investors, especially investment and pension funds, which are considered to react more rapidly to expected changes in inflation and output, have gained importance in the 1990s. But investments managed by Austrian investment and pension funds are still low by international comparison. However, growth is expected to accelerate within the next few years, as the state pension system is increasingly burdened.

Fourth, the Austrian stock market did not begin to develop until the end of the 1980s (i.e. share turnover totalled Sch 20 billion in 1989) and, even today, is marked by moderate liquidity concentrated in a few shares. Thus, it is not surprising that the stock price index does not provide much information.

Fifth, the information content of the yield curve might also have been affected by Austria's choice of exchange rate regime. Linking the schilling to the Deutsche mark generally implies a parallel movement of interest rates, but not necessarily of economic activity and inflation in the short term. Moreover, German reunification represents an exceptional period with possibly strong effects on the sample.

Looking ahead, the creation of the European Monetary Union will change the focus of research, also for the OeNB. Economic analysis will concentrate more strongly on the EMU area. At the OeNB, economic analysis will attribute great importance to financial indicators, in particular yield curve analysis. Yield spreads, for example, might be a good predictor of economic activity and inflation in the EMU, above all against the background of the significant results often found for larger economies.

2. Some remarks on the wealth transmission channel in Austria

In the context of the Austrian exchange rate and monetary policy of the last decades, attention – as already mentioned – was paid to the mutual dependencies of monetary policy measures, investment decisions, asset prices, and possible consequences for the goods markets only insofar as such measures should *not* influence the expectations of economic agents concerning the stability and credibility of the schilling/Deutsche mark peg. Disturbances, which otherwise would have imparted on the "real" economy, the wage-price link, and interest rates were regarded as potentially considerable and as endangering the concept itself. On the other hand, relatively higher interest rates which were sometimes necessary to defend the peg, were regarded as less harmful as their efforts were mitigated by interest rate subsidies, given for certain purposes, and by the high proportion of medium and long-term loans with fixed interest rates.

The view that effects of monetary policy on private assets and consequently on consumption and investment decisions were of minor importance may have its cause – at least as far as effects on consumers' activity are concerned – in the fact that wealth of relevant size in the form of shares, bonds and property etc. has accumulated only in the more recent past (if we disregard simple savings deposits). But also for the enterprise sector, decisions whether to invest in either real or financial capital were not very relevant for long periods after World War II.

A first attempt to quantify the importance of wealth effects within the transmission process of monetary policy was undertaken by Glück (1995) in connection with the BIS project on the monetary policy transmission process. The wealth effect was defined there as capturing the effects of monetary policy measures on the value of financial assets and consequently on consumption and investment. Because of data restrictions, "financial assets" had to be defined rather narrowly in this exercise (at least for Austria; it included deposits and an estimate for bonds in the portfolio of households – an approach quite frequently used as data problems are not unique to Austria). The outcome of the simulation experiment showed comparatively small effects for the wealth channel (compared, for instance, with the capital cost and the exchange rate channel). Furthermore it seemed to operate with rather long lags.

This was to be regarded, of course, as a quite global result and it would certainly be useful and sensible to improve the knowledge on further aspects and details of the wealth channel. How, for example, do policy measures affect asset prices and other wealth variables and what consequences does this have on aggregates like growth, employment, inflation, etc.? However, this intention is heavily impeded by the lack of reliable data. What can be done, therefore, is to investigate some additional pieces of evidence of interest in this context, try to find some tentative answers to the question as to which wealth effects and which effects on asset prices should be taken into account when formulating a possible monetary policy change.

Like the transmission process itself, the wealth channel can be viewed as a complex bundle of different effects. In the following, we will try to take a closer look at the effects of wealth changes on consumption, looking – as far as possible – at a broader spectrum of wealth; on investment decisions, not only as far as, for instance, the direct influence of higher interest rates on fixed capital investment is concerned, but also on portfolio decisions; on inflationary expectations; on banks' willingness to lend; and at the money demand function.

2.1 Consumption

The consumption channel of wealth effects has been strongly advocated, for example, in the MPS model. Based on Modigliani's life-cycle model, consumption spending is determined by the lifetime resources of consumers, which are made up of discounted income representing human capital, financial wealth and property. A major component of financial wealth is common stocks. Since a contractionary monetary policy can lead to a decline in stock prices, the value of financial wealth decreases, thus reducing the lifetime resources of consumers, and consumption should fall (Mishkin (1995)).

Formally, real consumption C is a function of life-cycle (permanent) income LCI and financial wealth FW:

$$C_t = f((LCI_t + FW_t)/P_t), \text{ with } P_t = \text{price index} \qquad (1)$$

Permanent income LCI can be determined with different degrees of complexity as the current and discounted future expected net income stream of households. The other determinant of private consumption, financial wealth FW, has been defined as (Dramais et al. (1997)):

$$FW_t = MV_t + F_t + B_t$$

i.e., it consists of the market value of firms in the domestic economy, MV, the net foreign asset position, F, and government debt, B. Though government debt enters the definition of private wealth, it has been discussed whether it has a positive effect on private consumption because households will deduct future tax payments and expected reductions in transfer payments, required to service the debt, from their permanent income. This proposition, known as Ricardian Equivalence, will, however, only hold in its extreme form of infinitely lived consumers. Life cycle consumers will discount the future more heavily and thereby underestimate the tax burden associated with government debt. Consequently, they regard government debt, at least partially, as net wealth. Some studies (e.g. Summers and Poterba (1987)) have shown, however, that this net wealth effect of government debt is negligible in the life-cycle model.

Some tests with the consumption equation used in Glück (1995) showed that government debt does not exert any significant influence, either on durable or on non-durable consumption. On the other hand, the influence of financial wealth, as used in this exercise, seems to have increased in the course of time, at least as far as durables are concerned, for which the elasticity with respect to wealth has risen from 1 in the period from the 1970s to the beginning of the 1980s to 2 in the period from the beginning of the 1980s to 1995.

The most difficult problem for estimations for Austria along the lines described at the beginning of this chapter is to find an appropriate wealth variable. In many studies we find a somewhat resigned attitude to this issue, for example in Gerdesmeier (1996): Wealth in a wide definition "... should include financial assets as well as real capital and human wealth. Quantifying the latter, however, faces unsurmountable problems. Consequently, it seems appropriate to rely on the sum of financial assets and real capital, the so-called non-human wealth" (p.10, our translation).

Others, e.g. Dramais et al. (1997), approximate human wealth as the discounted value of an infinite stream of labour income and transfers:

$$LCI_t = \int_t^{\infty} [(1 - tl)W_s N_s / P_s + TR_s / P_s] \exp - \int_t^{s} (r + p) dj \, ds$$

with LCI = life-cycle income, tl = tax on labour income, W = wage rate, N = number of people employed, P = price index, TR = transfers, r = interest rate, and p = probability of death.

We calculated a simplified form for use in equation (1).

For the second component in (1), financial wealth FW, an estimate of the market value of firms, MV, is required. Usually this is deducted from the firm's maximisation problem (see, e.g. Galeotti (1988)). In a first approximation we use the market capitalisation of firms quoted at the stock exchange.

Estimation of (1) yields disappointing results which are, in any case, inferior to the Brown-type consumption function augmented by the wealth term used in Glück (1995). Taking into account the findings of many other studies that a sizeable fraction of consumption is based on real current disposable income because of liquidity constraints and constructing a linear combination of the independent variable in (1) and current disposable income (as, e.g. in Dramais et al. (1997)), did not improve the results.

Digression on property

Meltzer (1995) emphasized that asset price effects extend beyond those operating through interest rates and equity prices. In his description of the Japanese experiences in the 1980s and 1990s he found that monetary policy had an important impact on the economy through its effect on land and property values. Generally, a monetary contraction can lead to a decline in income and property values, which causes households' wealth to shrink, thereby causing a reduction in consumption and aggregate output.

In econometric models, this property aspect of the wealth channel is frequently dealt with in relation to housing services, i.e. the interaction of demand and supply of housing services plays a decisive role for this aspect of the transmission of monetary policy to the household sector. In these approaches, as used, for instance, in the Bank of Finland model, the market price of housing is the discounted present value of the determinants of the rental price of housing and affects household sector wealth via the value of the housing stock and the accumulation of the housing stock. Monetary policy affects the market price of housing directly through the interest rate used in discounting. For the time being, lack of data does not allow the modelling of this aspect of the wealth channel for Austria.

2.2 Investment and portfolio decisions

In the context of investment and wealth effects, Tobin's q theory seems to be useful as it provides a mechanism through which monetary policy affects the economy by its effects on the valuation of equities. Tobin (1969) defines q as the market value of firms divided by the replacement cost of capital. If q is high, the market price of firms is high relative to the replacement costs, and new plant and equipment are cheap relative to the market value of business firms. Companies will then issue equity and get a high price relative to the cost of the plant and equipment they are buying. Investment spendig will rise as firms can buy much equipment for a small issue of equity (Mishkin (1995)). The influence of monetary policy on this process – apart from the role of the interest rate as a discounting factor in determining the market value of firms – is simply that monetary contractions raise the incentive for the public to hold more bonds and less stocks, thus reducing stock prices and the market value of the firms.

The q theory of investment has a number of theoretical advantages over competing models of investment. First, unlike the neoclassical model, it is forward-looking rather than being based on lags and past variables. Second, it allows for a distinct analysis of the effects of temporary versus permanent changes in tax parameters. Finally, it avoids the Lucas-critique, since the estimated adjustment parameters should not depend on policy rules (Schaller (1990)). Unfortunately, the theoretical appeal of the q theory has not been matched by empirical success. It is true that most of the studies find that investment is significantly related to q. However, in most cases variations in q explain only a small part of the variation in investment. The unexplained portion is usually highly serially correlated, and variables like profit and output, which should not matter according to the q theory, frequently exert a more significant influence on investment (Schaller (1990)).

Despite these problems and shortcomings we try to estimate an investment function based on the q theory as, to our knowledge, this has not been tried before for Austrian data.

It can be shown (Galeotti (1988)) that values for q can be approximated by:

$$q = MV/gK$$

with MV the aforementioned market value of firms, g the investment market price divided by the price of output, and K the capital stock.

The equations to be estimated can have the form:

$$I_t = a + bq_t + u_t \qquad \text{or} \tag{2}$$

$$(I_t/K_t) = a' + b'q_t + u'_t \tag{3}$$

with I = investment.

The first problem we encountered was the fact that because of the strongly rising values of MV our q does not oscillate around 1, as theory demands. We tried to overcome this by estimating on changes. The results then were in conformity with the ones mentioned above for other studies insofar as the explanatory power of (2) and (3) is very poor, as the significance of q, with one exception, never reaches the 5% level. The exception is the equation for construction for the period 1976-86, where the t-value for q reaches 2.32.

As far as portfolio decisions are concerned, Hahn (1990) investigated the potential influence of interest rate changes on portfolio decisions of large enterprises. He found that the portion of financial assets in relation to all assets correlated positively with the movements of bond yields (secundary market rate), i.e. the higher the bond yield, the higher the portion of assets held as financial wealth and the less invested in "productive" capital, such as machinery and equipment. Hahn also found evidence that large financial portfolios correlated negatively with rentability. He interpreted these results as an indication that low rentability and the weak investment performance of large enterprises could be caused by higher financial involvements, implying that large companies probably more often overlook profitable investment and innovation opportunities than smaller firms do.

In terms of the transmission mechanism, this would mean that raising interest rates would reduce investment, not only because of higher capital cost but also because of greater incentives to invest in financial assets, leading to a shift from real to financial assets. It can be supposed that high – and the perspective of further rising – asset prices would strengthen this tendency.

The simple but illustrative equation estimated by Hahn was:

$$FS_t = \underset{(9.799)}{3.117\, t} + \underset{(10.968)}{4.027\, RS_t}$$

$R^2 = 0.763$, D.W. = 1.605

with FS = financial asset as % of total assets, t = time, and RS = bond yield, secondary market. Figures in parentheses are t-values.

2.3 Price expectations

Monetary policy action changes price expectations, thus strongly influencing the discounted real values of permanent income, market valuation of firms etc. It is obvious that small variations of prices and interest rates change these discount factors heavily. For future work, some sensitivity analysis will be useful to check the magnitude of these effects.

2.4 Bank lending

In a recent study on the credit channel in Austria, Quehenberger (1997) found that the influence of monetary policy on credit conditions did not appear to have marked consequences for the investment activity of firms. Also, there is no evidence for credit rationing, as banks obviously try to refrain from quantity restrictions in tight monetary conditions. All in all, there does not seem to exist a credit channel in Austria as far as price and quantity of loans are concerned. It would need further investigation to see whether this result is modified when wealth effects are taken into account, i.e. whether changes in asset prices and market valuation via their influence on collateral affect the volume of loans. Some preliminary and very simple correlation and estimation attempts show that the market value of firms delivers a minor explanatory contribution in equations for the credit volume, not enough, however, for far-reaching conclusions.

2.5 Money demand

There is no room for monetary targeting in the concept of the Austrian exchange rate and monetary policy. This, however, will change soon, so that a more precise knowledge of the determinants of money demand will be necessary. When based on portfolio considerations, the influence of wealth on money demand has to be taken into account. The inclusion of wealth as an additional explanatory variable in money demand functions for Austria gives the correct sign, but wealth is insignificant. More detailed research, however, seems necessary.

Conclusion

This section of our paper documents our endeavour to push ahead the empirical knowledge on wealth effects within the transmission process of monetary policy in the Austrian economy. In this first attempt, we concentrated mainly on life-cycle approaches to consumption and Tobin's q theory of investment, which seemed most appropriate in this context. Unfortunately, our results are rather disappointing. This may have its cause, above all, in the very weak data base, with the consequence that in many cases we had to work with rough approximations. We hope that further research will bring some improvements.

Appendix 1: Data sources

IP: Industrial production at constant prices (seasonally adjusted monthly series; energy excluded). *Source*: WIFO (Austrian Institute for Economic Research).

INF: Annual change in the consumer price index (seasonally adjusted monthy series; all items), 1986=100.
Source: WIFO.

SM: (Sekundärmarktrendite) 10-year government bond yields (monthly averages).
Source: Oesterreichische Kontrollbank.

VIB: Interest rate on 3-month interbank loans (VIBOR) (monthly averages).
Source: OECD.

BMU: 10-year Treasury bond yield (benchmark bonds; monthly averages).
Source: Datastream.

BMB: 10-year government bond yield (benchmark bonds; monthly averages).
Source: Datastream.

B3M: Interest rate on 3-month interbank loans (monthly averages).
Source: Datastream.

Appendix 2

Estrella and Hardouvelis (1991)

Industrial production (cumulative growth rate)

Horizon (months)	Spread (SM-VIB)			
	Standard errors corrected with Newey-West procedure		Equation including AR(3) model for noise	
	Coefficient	t-statistic	Coefficient	t-statistic
6	1.37	1.40	0.34	0.29
12	1.86	2.07	-0.40	-0.47
18	2.36	2.91	-0.36	-0.60
24	2.32	3.06	0.52	1.03
36	1.72	3.23	-0.09	-0.24
48	1.55	4.93	0.38	1.43

Plosser and Rouwenhorst (1994)

Industrial production (cumulative growth rate)

Horizon (months)	Spread (SM-VIB)		VIB		Spread (SM-VIB)		VIB	
	Standard errors corrected with Newey-West procedure				Equation including AR(3) model for noise			
	Coefficient	t-statistic	Coefficient	t-statistic	Coefficient	t-statistic	Coefficient	t-statistic
6	-0.01	-0.84	-0.01	-2.21	-0.01	-1.20	-0.01	-1.69
12	-0.02	-1.90	-0.03	-4.92	-0.03	-2.72	-0.03	-3.44
18	-0.03	-2.47	-0.04	-6.77	-0.03	-2.70	-0.04	-3.76
24	-0.04	-2.84	-0.06	-6.67	-0.02	-1.69	-0.04	-4.21
36	-0.08	-3.61	-0.08	-7.32	-0.03	-1.89	-0.05	-2.91
48	-0.02	-0.57	-0.05	-2.70	0.00	0.25	-0.02	-0.94

Fama (1990)

Changes in inflation

Horizon (months)	Spread (SM-VIB)		Inflation		Spread (SM-VIB)		Inflation	
	Standard errors corrected with Newey-West procedure				Equation including AR(1) model for noise			
	Coefficient	t-statistic	Coefficient	t-statistic	Coefficient	t-statistic	Coefficient	t-statistic
6	0.00	-3.28	-0.35	-3.92	0.00	-0.58	0.00	0.07
12	-0.06	-3.43	-0.71	-7.64	0.00	-0.33	0.08	0.60
18	-0.01	-3.38	-0.98	-10.23	0.00	-0.54	-0.14	-1.23
24	0.00	-2.46	-1.18	-10.96	0.00	-0.38	-0.01	-0.12
36	0.00	0.81	-1.33	-13.51	0.00	0.60	-0.13	-1.28
48	0.01	2.76	-1.30	-23.82	0.00	0.11	0.05	0.42

Appendix 3: Vector error correction estimates

Cointegrating Eq1					
Log(*IP*(-1))	1.000000	Log(*INF*(-1))	-0.100213 (0.03158) (-3.17323)	(*SM*(-1)–*VIB*(-1))	-0.059646 (0.01457) (-4.10849)
(*BMB*(-1)–*BMU*(-1))	-0.072688 (0.01072) (-6.78089)	(*SM*(-1)–*BMB*(-1))	-0.040121 (0.02704) (-1.48399)	(*BMB*(-1)–*B3M*(-1))	-0.000437 (0.01552) (-0.02813)
C	-4.474237 (0.04059) (-110.227)				

Error correction	*D*(log(*IP*))	*D*(log(*INF*))	*D*(*SM-VIB*)	*D*(*BMB-BMU*)	*D*(*SM-BMB*)	*D*(*BMB-B3M*)
CointEq1	0.034030 (0.04481) (0.75942)	0.695906 (0.26027) (2.67379)	0.109866 (0.37036) (0.29665)	1.683871 (0.44436) (3.78941)	0.362445 (0.28776) (1.25952)	-0.834282 (0.46287) (-1.80242)
D(log(*IP*(-1)))	-0.631555 (0.09025) (-6.99807)	-1.208133 (0.52418) (-2.30480)	0.533311 (0.74590) (0.71499)	-0.588403 (0.89494) (-0.65748)	-0.234749 (0.57955) (-0.40505)	0.103271 (0.93221) (0.11078)
D(log(*IP*(-2)))	-0.369388 (0.08384) (-4.40597)	-0.919513 (0.48696) (-1.88829)	0.663936 (0.69293) (0.95816)	-0.899932 (0.83139) (-1.08245)	-0.917188 (0.53840) (-1.70356)	1.309361 (0.86601) (1.51195)
D(log(*INF*(-1)))	-0.000677 (0.01443) (-0.04694)	0.029589 (0.08381) (0.35302)	0.132601 (0.11927) (1.11180)	-0.291888 (0.14310) (-2.03977)	0.147253 (0.09267) (1.58903)	-0.034012 (0.14906) (-0.22818)
D(log(*INF*(-2)))	0.024908 (0.01462) (1.70380)	0.004373 (0.08491) (0.05150)	-0.151528 (0.12083) (-1.25408)	0.064512 (0.14497) (0.44499)	1.159683 (0.09388) (1.70088)	-0.044125 (0.15101) (-0.29220)
D(*SM*(-1)-*VIB*(-1))	0.004865 (0.01116) (0.43577)	-0.052293 (0.06485) (-0.80637)	0.127137 (0.09228) (1.37775)	0.083590 (0.11072) (0.75497)	0.025188 (0.07170) (0.35130)	0.077093 (0.11533) (0.66846)
D(*SM*(-2)-*VIB*(-2))	0.008367 (0.01058) (0.79051)	-0.008058 (0.06147) (-0.13108)	-0.150626 (0.08748) (-1.72192)	-0.030485 (0.10496) (-0.29046)	-0.029205 (0.06797) (-0.42969)	-0.026892 (0.10933) (-0.24598)
D(*BMB*(-1)-*BMU*(-1))	0.014429 (0.00768) (1.87876)	-0.038674 (0.04461) (-0.86697)	-0.217930 (0.06348) (-3.43327)	0.395090 (0.07616) (5.18766)	0.163437 (0.04932) (3.31380)	-0.196282 (0.07933) (-2.47421)
D(*BMB*(-2)-*BMU*(-2))	-0.002627 (0.00822) (-0.31951)	-0.019564 (0.04775) (-0.40968)	0.055760 (0.06795) (0.82058)	-0.289134 (0.08153) (-3.54631)	0.050833 (0.05280) (0.96277)	0.019755 (0.08493) (0.23262)
D(*SM*(-1)-*BMB*(-1))	-0.008102 (0.01364) (-0.59421)	-0.030063 (0.07920) (-0.37958)	-0.093421 (0.11270) (-0.82894)	0.085602 (0.13522) (0.63306)	0.196798 (0.08757) (2.27026)	0.016940 (0.14085) (0.12027)
D(*SM*(-2)-*BMB*(-2))	0.008969 (0.01341) (0.66892)	-0.171114 (0.07788 (-2.19728)	0.060680 (0.11061) (0.54758)	0.183520 (0.13296) (1.38028)	-0.203299 (0.08610) (-2.36115)	0.194368 (0.13849) (1.40344)
D(*BMB*(-1)-*B3M*(-1))	0.000832 (0.00967) (0.08608)	0.043755 (0.05614) (0.77942)	0.178181 (0.07968) (2.23052)	0.015660 (0.09585) (0.16339)	-0.033615 (0.06207) (-0.54158)	0.216755 (0.09984) (2.17109)
D(*BMB*(-2)-*B3M*(-2))	0.001316 (0.00976) (0.13475)	0.046853 (0.05672) (0.82609)	0.117595 (0.08071) (1.45707)	0.081289 (0.09683) (0.83948)	0.043397 (0.06271) (0.69205)	-0.033840 (0.10087) (-0.33550)
R^2	0.332448	0.136271	0.211907	0.290520	0.211349	0.159237
Adjusted R^2	0.275229	0.062237	0.144356	0.229707	0.143750	0.087172
Sum square residuals	0.075950	2.562267	5.188230	7.468846	3.132204	8.103821
S.E. equation	0.023292	0.135285	0.192507	0.230974	0.149576	0.240592
Log likelihood	346.9237	95.75318	41.78241	13.90979	80.38857	7.667764
Akaike AIC	-7.438179	-3.919611	-3.214111	-2.849763	-3.718766	-2.768168
Schwarz SC	-7.180691	-3.662123	-2.956622	-2.592275	-3.461278	-2.510679
Mean dependent	0.002715	-0.004346	0.017974	0.022477	-0.008301	0.001170
S.D. dependent	0.027359	0.139702	0.208113	0.263169	0.161644	0.251818

Determinant residual covariance	8.78E-12	Akaike information criteria	-25.27330
Log likelihood	1,103.969	Schwarz criteria	-24.99271

Notes: Sample (adjusted): 1983:04 – 1995:12. Included observations: 153 after adjusting end-points. Standard errors and t-statistics in parentheses.

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Asset prices: relationships with demand factors and credit, and implications for monetary policy

Pierre Jaillet and Pierre Sicsic[1]

Introduction

The middle of the last decade was marked in France by financial deregulation, and the years that followed saw a sharp increase in the price of equities and property in Paris, a substantial decline in the household saving rate and an acceleration in corporate investment. This chronological account might suggest that pronounced wealth effects are at play in France, and that there may be a link between financial deregulation and rising asset prices, probably through an expansion of lending to asset buyers.

The aim of this paper is to compare these theories with the available data. The latter do not validate the hypothesis of a wealth effect. It does not appear possible to show a statistical relationship between household consumption and equity prices; the rise in residential property prices may not have appreciably affected the whole of the country but appears to have been confined to Paris; the increase in lending has been markedly greater than that in productive investment (firms' gross fixed capital formation), and this development may have contributed to the rise in commercial property prices.

Even if there is no direct relationship between asset prices and demand factors, movements in asset prices may have an effect on the transmission of monetary policy if, by changing banks' balance-sheet positions, they affect the relationships between market rates and lending rates. It is conceivable that, following a deterioration in their balance-sheet strength linked to a downturn in the value of the collateral for their loans, banks will pass on reductions in market rates less widely. Contrary to this theory, the adjustment of lending rates to market rates has become less and less sluggish in recent years.

1. Equity holdings of households[2]

Stock market capitalisation in France stood at FF 3,073 billion in December 1996 and FF 2,415 billion at the end of 1994, or 33% of GDP. Capitalisation is lower than in the United States, where it amounted to $6,049 billion at the end of 1994, or 87% of GDP. Moreover, the share of capitalisation accounted for by households is much smaller in France: whereas 64% of quoted shares were controlled by households in the United States (including pension funds' holdings of equities), the ratio was at most 30% in France (including share portfolios of UCITS held by households).[3]

Not very much is known about the equity holdings of French households. At the end of 1994, according to the Bank of France's securities survey, listed French shares (and share-based UCITS) held by households amounted to FF 443.4 billion (of which FF 128.7 billion in share-based UCITS). As the overall coverage of this survey was 61.4% of stock market capitalisation, households'

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[1] Members of the Economics Department.

[2] We benefited from the help of Gunther Capelle-Blancard in drafting this section.

[3] The data on the United States are from Poterba and Samwick (1995), p. 323.

capital in listed shares can be put at a little over FF 700 billion, or approximately 30% of stock market capitalisation.[4] Total holdings of equities by French households are therefore much lower than those of US households.

With regard to the concentration of equity holdings, in the United States the 10% of households with the largest equity portfolios accounted for 90% of total holdings of equities by households; the 12% of US households with the highest income held 58% of the total equity stock, including pension funds, in 1992 (Poterba and Samwick, 1995, Tables 9 and 10). In France, according to INSEE's financial assets survey of 1992, the wealthiest 10% of households (in terms of total financial wealth, ranging from equities to cheque accounts) accounted for 86% of total securities held by French households (Arrondel, 1996, Table 9C, p. 54). It is likely that the concentration is even greater for shares alone. Taking income and total financial wealth as equivalent criteria for classifying households for purposes of international comparison, it appears that the concentration of equity holding is greater in France than in the United States.

Taking account of the concentration of equity holdings, a wealth effect from a rise in equity prices should be reflected in a relative increase in the consumption of products purchased by the wealthiest households. Such a correlation does not appear in the United States. Poterba and Samwick (1995, p. 297) do not find any wealth effect from equity prices on household consumption and conclude that the positive correlation between equity prices and total consumption is due to the fact that price movements act as a leading indicator. The issue of the correlation between equity prices and household consumption in France is addressed in the following section.

2. Equity prices and household consumption[5]

The rise in equity prices on the Paris Stock Exchange was particularly sharp from 1985 to 1990. It can be seen from Figure 1 that stock market prices increased more rapidly in Paris than in New York, London or Frankfurt over this period. At the same time, the saving rate of French households declined considerably, which suggests that there is a statistical relationship between equity prices and household consumption.

The first run to test the relationship between the saving rate and the boom in the stock exchange is to regress the quarterly growth rate of household consumption on past real growth rates of the French equity index. From Table 1, it can be seen that the (almost) significant effect of the increase in equity prices disappears when the dummy for the last quarter of 1974 is introduced.[6]

When income is introduced to estimate the usual ECM consumption *à la* Hendry, the coefficient of the increase in the real stock index remains insignificant.[7]

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4 443.4 billion divided by 0.614 gives 722.1 billion. This information is from Chocron and Marchand (1995, pp. 161 and 165). This estimate overstates households' equity wealth as the coverage of the securities survey, conducted among banking establishments which manage securities accounts, was probably higher for households than for other agents. According to the quarterly financial operations tables, households held FF 779 billion of listed French shares (excluding share-based UCITS) at the end of 1994. Arrondel et al. (1996, p. 158) present the various estimates of the equity wealth of French households and prefer the total calculated on the basis of the replies to the Bank of France's securities survey, which is close to that calculated using the replies to the financial assets survey conducted among households by INSEE.

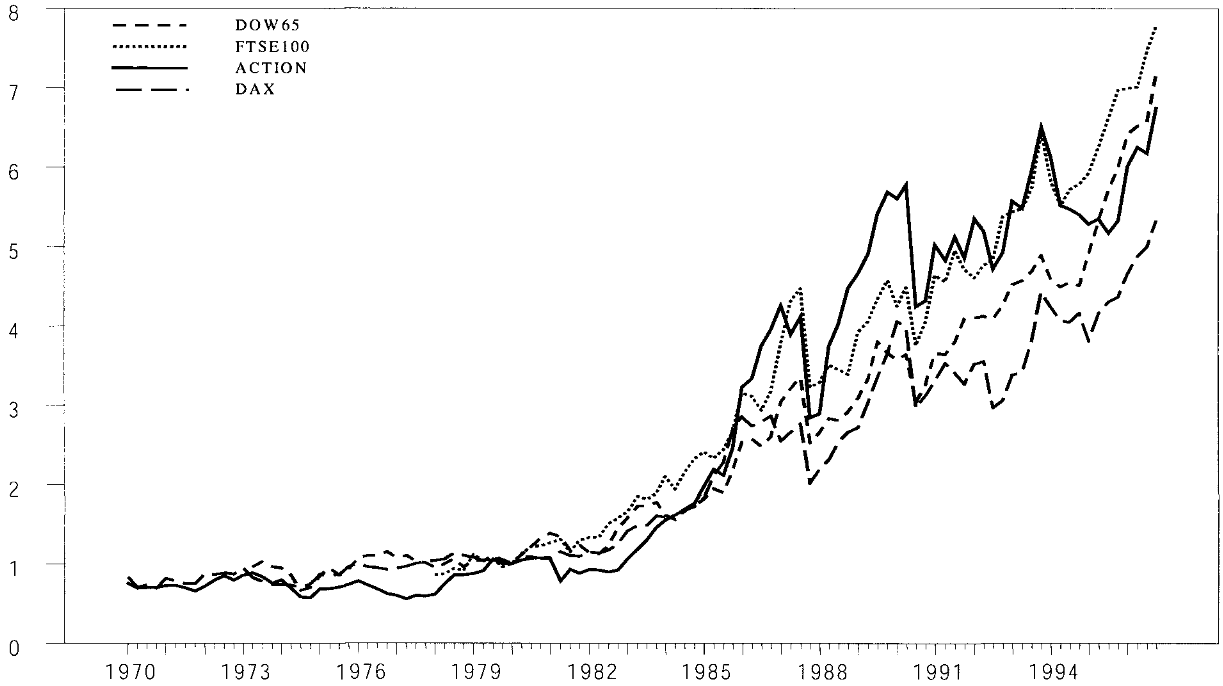
5 The regressions in this section were carried out by Jean-Pierre Villetelle.

6 This dummy is introduced since measured consumption is an outlier owing to strikes.

7 This equation does not model consumption as a function of financial wealth since the stock index is not included in the correction term.

Figure 1

Nominal stock exchange indices



Note: All the indexes have been rebased to 1 in 1980:Q1. ACTION is the SBF250 index after 1990:Q4. From 1987:Q3 to 1990:Q4 it follows the growth of CAC40. Before 1987:Q3 it follows the index for French assets with variable earnings published in the *Bulletin mensuel de statistique*.

Table 1

Estimations from 1974:1 to 1995:4

	E0		E1		E2	
	Coeff.	**t-stat.**	**Coeff.**	**t-stat.**	**Coeff.**	**t-stat.**
a_0	0.0057	7.3	0.0056	7.2	0.0058	7.6
w_0	-0.0068	-0.9				
w_1	0.0083	1.1				
w_2	0.0150	2.0	0.0139	1.9	0.0110	1.5
w_3	-0.0048	-0.6				
w_4	-0.0083	-1.1				
k_0					-0.015	-2.1
DW	2.1		2.1		2.1	
SER (%)	0.73		0.73		0.71	
R^2	0.42		0.40		0.43	

Regression of quarterly consumption growth ($\Delta\log C$) on current and lagged values of real growth of stock index ($\Delta\log Icact$), given by:

(E0): $\Delta \log C = a_0 + w_0 \Delta \log Icact + w_1 \Delta \log Icact_{-1} + w_2 \Delta \log Icact_{-2} + w_3 \Delta \log Icact_{-3} + w_4 \Delta \log Icact_{-4}$

(E1): $\Delta \log C = a_0 + w_2 \Delta \log Icact_{-2}$, which is equation (E0) after suppression of the non-significant terms.

(E2): $\Delta \log C = a_0 + w_2 \Delta \log Icact_{-2} + k_0 D_{74:4}$, which is equation (E1) with a dummy variable in 1974:4.

An equation which includes a dummy variable equal to 1 from 1986 to 1990 (Ea3, in Table 2), shows that the saving rate was abnormally low in relation to income and inflation during this period. This episode is not linked to equity prices (the coefficient of the increase in the real share price index is very far from being significant) but to the expansion in consumer credit, which was one important aspect of the financial liberalisation of this period.

These regressions show that neither the crude regression of consumption growth on stock index growth, nor specifications of consumption in terms of income and variables taking account of the financial deregulation of the mid-1980s (discussed in Sicsic and Villetelle (1995)) indicate any significant influence of stock prices on households' consumption in France.

Table 2

Estimations from 1974:1 to 1995:4

	Ea		**Ea with 1974:4 dummy**		**Eac**		**Ea3**	
	Coeff.	**t-stat**	**Coeff.**	**t-stat**	**Coeff.**	**t-stat**	**Coeff.**	**t-stat**
a	-0.11	-4.7	-0.11	-5.5	-0.15	-12.3	-0.15	-16.1
b	-8.58	-2.8	-8.10	-3.1	-5.81	-4.3	-4.64	-4.7
c	-0.09	-2.8	-0.10	-3.2	-0.15	-4.6	-0.17	-5.0
d	-2.55	-3.3	-2.41	-3.5	-2.21	-5.0	-1.81	-4.7
w_2	0.0134	1.9	0.0106	1.5	0.0066	1.0	0.0041	0.6
$-w_2/c$	0.15		0.11		0.04		0.02	
e					3.55	3.7	1.46	1.7
100 u							3.34	2.7
k			-0.016	-2.3			-0.017	-2.7
DW	2.26		2.27		2.40		2.48	
SER %	0.67		0.65		0.63		0.60	
R^2	0.50		0.53		0.56		0.62	

Regressions of consumption on income (R) and inflation ($\Delta\log p$), augmented by the growth rate of the stock index ($\Delta\log Icact$), such that:

(Ea): $\Delta \log C = -ac + (1-cb)\Delta \log R + c \log \frac{C_{-1}}{R_{-1}} - dc\Delta \log p + w_2 \Delta \log Icact_{-2}$

(Eac): $\Delta \log C = -ac + (1-cb)\Delta \log R + c \log \frac{C_{-1}}{R_{-1}} - dc\Delta \log p + w_2 \Delta \log Icact_{-2} + e(1-cb)\frac{\Delta CRT}{pR} + e(-1+cb-c)\left(\frac{\Delta CRT_{-1}}{p_{-1}R_{-1}}\right)$,

which is equation (Ea) augmented by consumer credit variation (ΔCRT).

(Ea3): $\Delta \log C = -ac + (1-cb)\Delta \log R + c \log \frac{C_{-1}}{R_{-1}} - dc\Delta \log p + w_2 \Delta \log Icact_{-2} + e(1-cb)\frac{\Delta CRT}{pR} + e(-1+cb-c)\left(\frac{\Delta CRT_{-1}}{p_{-1}R_{-1}}\right)$

$-cuD_{86-90} - ckD_{74:4}$,

which is equation (Eac) augmented by a dummy variable for 1986:1-1990:4 and by a dummy variable in 1974:4.

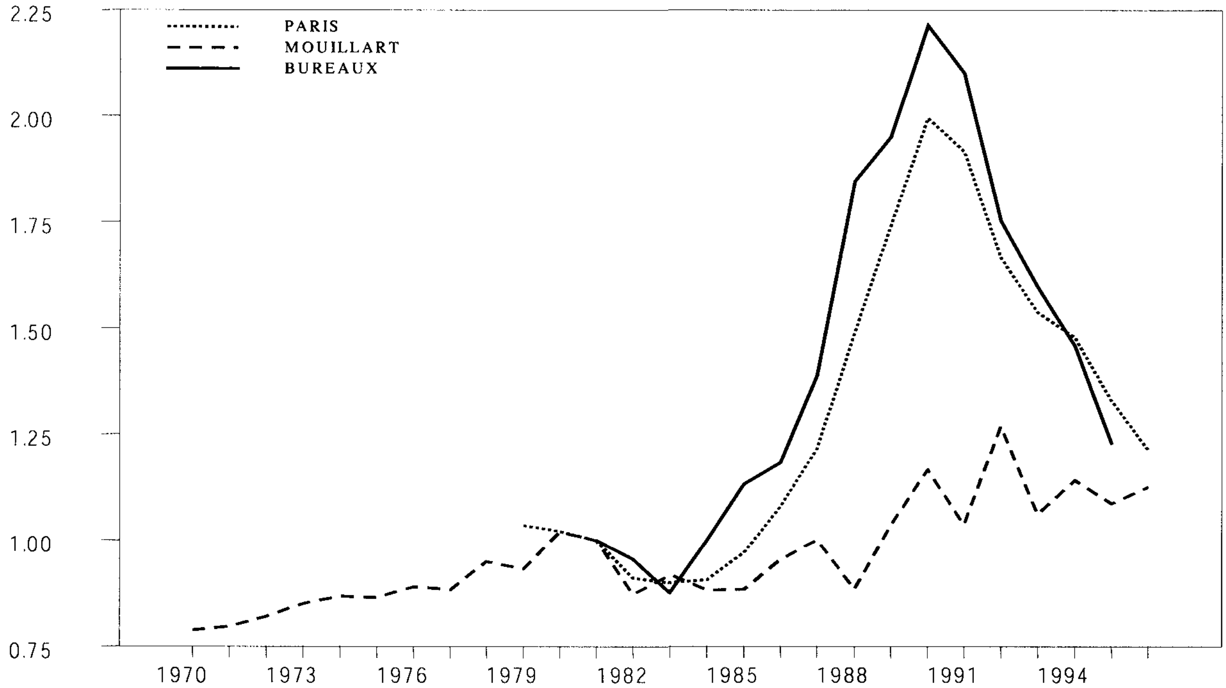
3. Lending and property prices

Property prices in Paris increased sharply at the end of the 1980s. The average price per square meter, according to the Paris Chamber of Notaries, doubled from 1985 to 1990, after account is taken of the rise in consumer prices. However, in general, residential property prices in France followed the pattern of inflation according to the Mouillart series (1997).[8]

Commercial property prices were calculated on the basis of data supplied by the Direction Générale des Impôts on average transaction values for each Paris arrondissement and for some other cities.[9] For each locality, two types of transaction values are available: for "luxury" offices and for "standard" offices. Four series were calculated by combining the locality (Paris/provinces) and the type of office (weighted according to the number of transfers of ownership), and then a simple average of these four series (divided by the consumption deflator and rebased to 1 in 1981) represents the real value of commercial property in France. The commercial property prices in Paris tripled from 1981 to 1990 and were, in 1987 and in 1995, 40% above their 1981 level.

Figure 2

Inflation-adjusted property price indices



Note: PARIS is the mean price of housing transaction according to the Paris Chamber of Notaries; MOUILLART is the index of housing for France kindly provided by Mouillart (1997); BUREAUX is the computed index for office building from data provided by the DGI. All indexes have been divided by the consumption deflator, and rebased to 1 in 1981.

The moderate and steady rise in the real price index for residential property for the whole of France (13% from 1981 to 1996; see Figure 2) shows that no wealth effects can be sought there. On the other hand, the strong increase observed in commercial property prices between 1985 and 1990, precisely at the time of financial deregulation, raises the question of whether there may be a relationship between credit expansion and rising asset prices.[10]

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8 This series will be taken up in the forthcoming national wealth accounts base (Moreau (1997)).

9 Thierry Grunspan kindly provided us with the unpublished data he obtained from the Direction Générale des Impôts, data which he has independently used in Grunspan (1997).

10 This is the basic idea advanced in the case of France, in particular by Borio, Kennedy and Prowse (1994).

Table 3

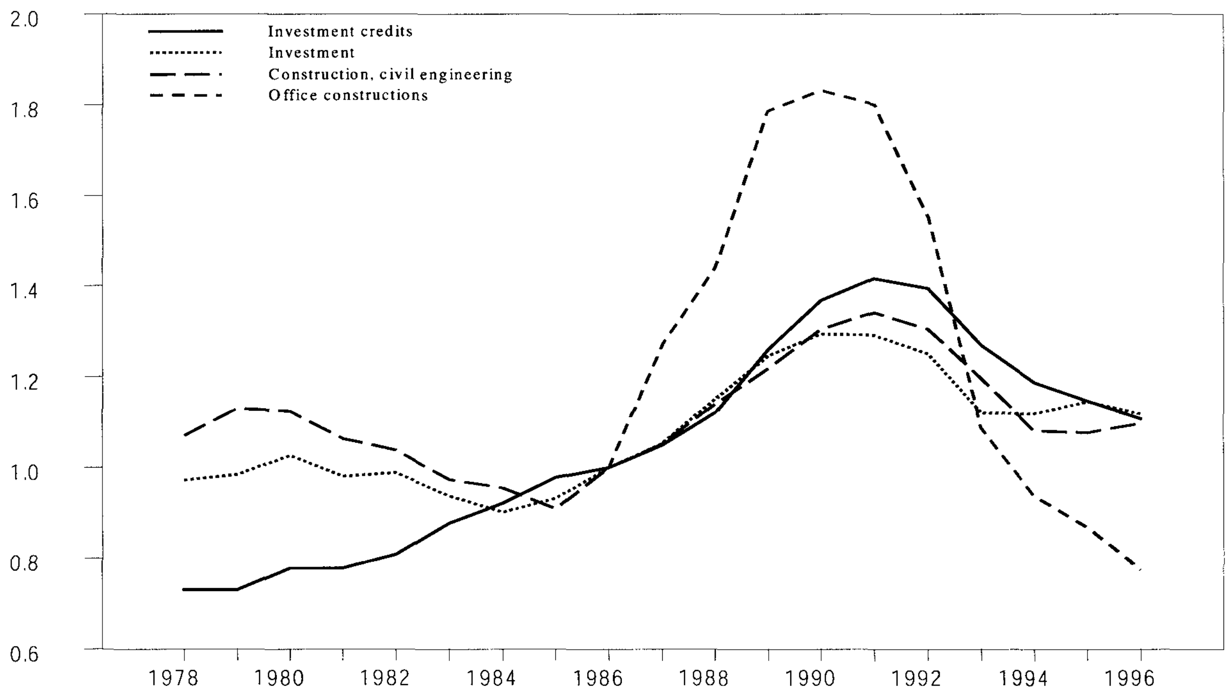
Average annual growth rates and contributions to the growth of total bank lending

In percentages

	1978–85		1985–91		1991–96	
	Growth rate	Contribution	Growth rate	Contribution	Growth rate	Contribution
Cash facilities for firms	10.18	1.64	10.45	1.46	-0.68	-0.09
Exports	5.15	0.26	-15.89	-0.24	-12.31	-0.05
Investment by firms	14.87	2.67	9.53	1.86	-2.90	-0.52
Investment by other agents	11.40	2.00	6.79	1.12	1.62	0.23
Consumer credits	17.59	0.54	18.37	0.94	4.63	0.29
Housing credits (home buyers)	13.71	4.94	7.02	2.50	2.06	0.66
Housing credits (developers)	-2.51	-0.01	49.58	0.52	-18.99	-0.36
Other lending	7.79	0.23	22.45	0.98	-12.39	-0.55
Doubtful debt	32.78	0.37	24.92	0.64	12.67	0.64
Total	12.61		9.75		0.22	

Figure 3

Investment credits, investment and construction indices



Note: Investment credits is an index of investment credits to firms; Investment is the gross capital formation of enterprises; Construction, civil engineering. is the gross capital formation of enterprises in construction and civil engineering; Office constructions is the number of new office constructions. The three first series have been divided by the consumption deflator.

Lending grew sharply between 1985 and 1991 (by 9.8%, compared with 12.6% in 1978-85; average inflation rates were 2.5% after 8.9%). Property development credit increased on average by almost 50% a year between 1985 and 1991, with the bulk of the expansion coming between 1988 and 1991. The total of such property development credit has always been limited (FF 175 billion in 1991, whereas other housing credits amounted to FF 1,996 billion). It should be emphasised that loans for financing residential or commercial property transactions carried out, for example, by property managers are classified, for the purpose of the monetary statistics, as corporate investment credits. These credits increased by 9.5% per annum from 1985 to 1991, more rapidly in real terms than from 1978 to 1985 (7%, compared with 6%).

Figure 3 shows that the increase in investment credits from the monetary statistics has been larger than the increase in the gross fixed capital formation by enterprises. With the self-financing ratio improving, the amount to be financed by external funds has actually declined since 1982. The increase in "investment" credits has obviously permitted other types of transactions than gross fixed capital formation. The number of new office constructions has, interestingly, followed the price of commercial property.

The stance of monetary policy was not accommodating in the mid-1980s, judging by the level of real interest rates. In view of the expected rates of return on asset holdings, the prices of which were rising sharply, the demand for credit is not very sensitive to interest rates, which are bound to be a good deal lower than these rates of return. The most famous historical example is that of the stock market boom in the United States between 1926 and 1929 which, according to White (1990, pp. 159 and 164), developed under a tight monetary policy and attracted funds through the reallocation of credit on the financial markets.

4. Transmission of short-term market rates to lending rates[11]

A potentially dangerous effect of asset price inflation, and subsequent deflation, stems from the deterioration of banks' balance sheets, which would translate into a stiffening of monetary policy transmission mechanisms, i.e. greater inertia in the lending rates applied by banks to borrowers in relation to market rates.

To test this idea, we used the individual data from the surveys conducted by the Bank of France among bank branches to measure the cost of credit. Figures are available for the lending rate and loan size. An econometric equation for each type of credit, controlling for the loan size effect and the fixed effects relating to type of bank and area of activity, was estimated up to 1994 to take account of an initial period of falling market rates, and projected thereafter.

The estimated equation is:

$$r_{k,t} = c + \sum_{i=1}^{3} \alpha_i tdir_t + \mu_1 C_k^1 + \mu_2 C_k^2 + \gamma_1 (tdir_t - \overline{tdir}) C_k^1 + \gamma_2 (tdir_t - \overline{tdir}) C_k^2 + \sum \kappa_i I_{i,k} + u_{k,t}$$

where $r_{k,t}$ is the lending rate for observation k in month t (the survey is carried out in January, April, July and October), $tdir_t^i$ is the three-month PIBOR in $t - i$, C_k^1 and C_k^2 are dummy variables for loan size, and the I_k^j are dummy variables for fixed effects of industry sector and bank sector. The simulated lending rates (not shown) are on the whole higher than the lending rates observed, which indicates that the passthrough of the fall in market rates to lending rates did not slow down.

Another way of describing this phenomenon is to show the market rate coefficients obtained in the regressions explaining the lending rates over an extended estimation period. These

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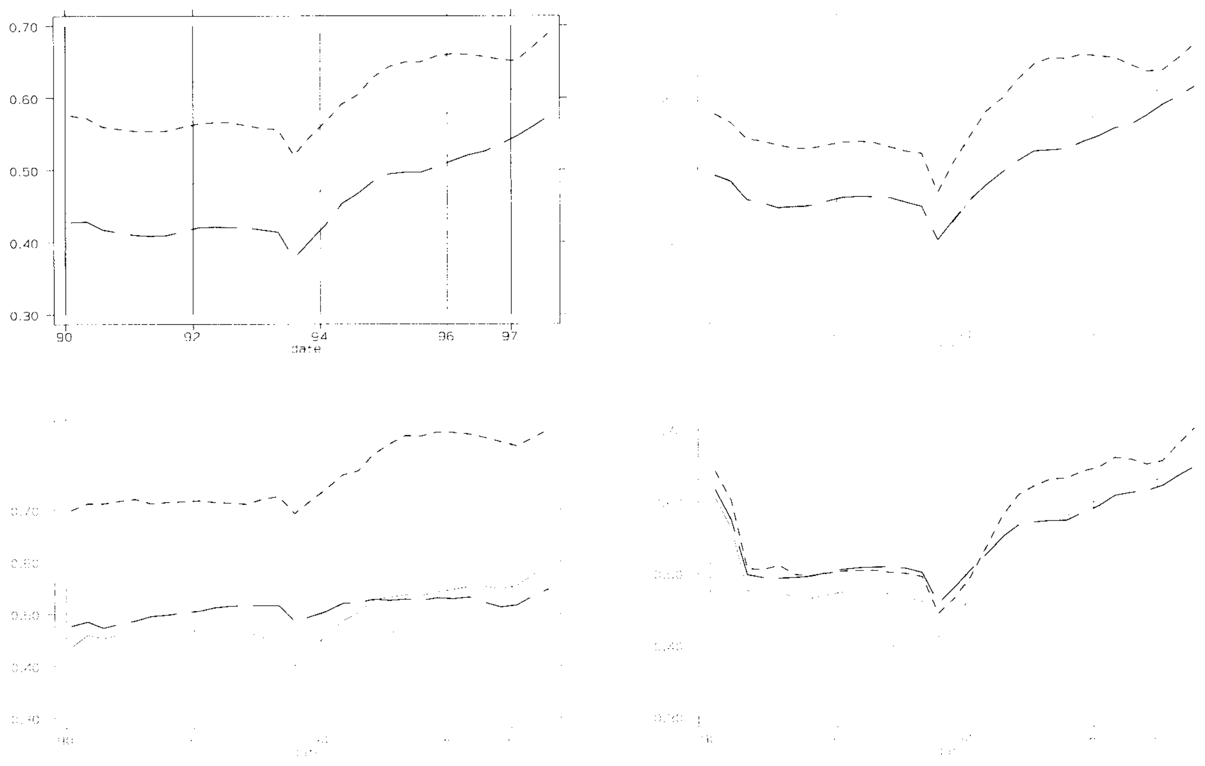
[11] Renaud Lacroix contributed to this section, producing the econometric estimations using the model proposed by Rosenwald (1998).

coefficients are $\sum_{i=1}^{3} \hat{\alpha}_i + \hat{\gamma}_1 + \hat{\gamma}_2$ for loans size 2 (see Figure 4). These coefficients might be expected to decrease when the sample incorporates periods of low market rates, as then lending rates come close to the break-even point beyond which fixed costs make it difficult for banks to cut them further. In line with intuition, the coefficients are higher for the largest amounts. The striking result, however, is that the coefficients linking lending rates to market rates have recently increased, while the (not shown) constant coefficients (the $\hat{\mu}_0 + \hat{\mu}_1 + \hat{\mu}_2$ for instance) went down. Lending rates have therefore not been particularly inert in the face of the fall in market rates in 1995-97.

Figure 4

Recursive estimators of the impact of market rates on lending rates



Note: Dotted line for the largest size loans, dashed lines for the smallest size loans. Upper graphs for discounts (left) and overdrafts (right), bottom graphs for long-term (left) and short-term (right) loans.

5. Should the monetary authorities react to asset price movements?

Most central banks today set themselves a final objective of *price stability*, defined on the basis of current inflation indicators (consumer price indices or national income deflators). The decoupling observed in many countries during the second half of the 1980s between the usual price indicators and asset prices (financial or property), however marked it may have been, has, fundamentally, changed nothing in this respect, despite the fact that, in their communications with the markets and the general public, the monetary authorities today appear more attentive to developments in asset markets (cf. Mr. Greenspan's statements on the "irrational exuberance" of the stock market, the references to the yield curve in the presentation of monetary objectives in France or in Germany, etc.).

This relative indifference of the monetary authorities to the, at times, erratic movements of asset prices can be justified in theory: assuming efficient markets, the indices of goods and services prices, provided that they are correctly weighted and reflect the whole range of goods and services prices, must embody all the information available on the future course of asset prices. Conversely, the latter must reliably foreshadow the price of the related services: a change in property prices, for example, must faithfully reflect the present values of expected rents over the whole life of the asset. From that standpoint, there are consequently no grounds for doubting the validity of the current inflation index, and hence of the nominal anchor used by central banks.

The problem is that the decoupling of indicators of goods and services prices from asset prices may temporarily assume such proportions as to lead to a rejection of the assumption of market efficiency and the homogeneity of price indicators; it must then be admitted that the formation of "speculative bubbles" on certain asset markets reflects a dynamic inherent in these markets, disruptions in which may distort the allocation of resources, amplify the real cycle via the debt channel and finally jam the transmission mechanisms of monetary policy to such an extent as to dangerously reduce its effectiveness. Confronted with such a scenario, a central bank naturally cannot maintain an attitude of benign neglect but must answer three sets of questions:

(i) *What are the available and reliable indicators for movements in asset prices* and in what respect can they cast new light on demand, activity and the maintenance of price stability? First it may be observed that the statistical information is generally of poor quality, in particular as far as real estate is concerned. Its partial character makes it rather inappropriate for constructing composite indicators.[12] This may at least partly explain the dearth of empirical work on the topic, its rather heterogeneous character, the uncertain or inconclusive nature of its results and, perhaps even more so, the difficulty of making international comparisons. As part of their main mission of maintaining price stability, the central banks must therefore remedy the current deficiency of statistical information.

(ii) *In what respect can monetary policy be held responsible for the temporary disruptions observed on the asset markets?* Borio et al. (1994), for example, reveal that credit contributed significantly to the sharp rise in asset prices as from the mid-1980s in a number of countries, including France. In the case of France, this is related to the introduction of a new monetary control regime (lifting credit ceilings and adoption of indirect procedures of liquidity management through the control of interest rates), without, however, it being possible to infer any causality from this: econometric tests would probably not confirm any overall relationship between lending and asset prices as the total of outstanding loans has stagnated since 1991 while equity prices have continued to rise. Statistics, nevertheless, suggest a relationship between certain categories of credit and the boom in prices on specific asset markets. These observations illustrate that, in the context of a monetary policy of indirect control via interest rates, central banks naturally lack the instruments needed to counter a disruption occurring on a specific market, such as the commercial property market. A matter of concern is that if this disruption results in a serious deterioration in the balance sheets of financial intermediaries, it makes them a priori less likely to react to monetary policy impulses. In the case of France, however, the econometric test results (see above) do not seem to validate this perception, as the setting of lending rates by credit institutions has not been significantly affected overall.

To conclude on this point, it is clear that this is where monetary and prudential policies intersect. The prudential authorities today appear more equal to the task of preventing excessive sectoral risks; in this respect, the close links that they may maintain with the central banks and

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12 The Bank of France is developing a composite asset price index in order to test the validity of a specific monetary policy transmission channel (Grunspan (1997)).

the common view they share cannot but foster the convergence of these microeconomic concerns and the objective of long-term monetary stability.

(iii) *What lessons can the central banks draw from this when formulating their targeting policies?* There can obviously be no question of defining a target in asset price terms. Apart from the problem of the availability and selection of a relevant indicator, a central bank adopting such a strategy would be confronted with two major difficulties: on the one hand, there is no reason why the lags with which asset prices and the prices of goods and services are affected, just as the necessary mix of interest rate changes, should be identical; thus, to effectively counteract the emergence of a speculative bubble, the central bank could be driven to operate large scale (and perhaps pro-cyclical) moves of its interest rate at the risk of seriously affecting activity, or endangering its final objective of price stability. On the other hand, it would be very difficult for a central bank to explain to the markets and the public why it had to tighten its interest rate policy in response to a rise in asset prices at the very time when all the usual inflation indicators were converging towards disinflation, or even below the upper bound of its target range, if any.

Thus, failing a change in the formulation of monetary policy objectives, the priority should be given to improving our knowledge of the links between asset prices and the financial indicators monitored by the central banks with a view to further reinforcing the preventive character of monetary policies.

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Asset market hangovers and economic growth: US housing markets

Matthew Higgins and Carol Osler*

Asset market bubbles matter to policymakers. For example, in December 1996, Alan Greenspan, Chairman of the Board of Governors of the Federal Reserve System of the United States, asked publicly: "How do we know when irrational exuberance has unduly escalated asset values, which then become subject to unexpected and prolonged contractions ...?" Stock market participants interpreted this comment – correctly or not – as a warning that stock prices might be overvalued. The market suffered a brief reversal, but bounced back and was soon reaching new highs. In February 1997, Greenspan used his testimony to the US Senate Banking Committee to cite the possibility of "excessive optimism" in the stock market.

This paper addresses one reason for policymakers' concern about asset market bubbles: bubbles can adversely affect real activity as they collapse. We estimate models of house prices and investment for US state housing markets, and arrive at two main results. First, house prices may be subject to speculative bubbles. Second, housing investment responds noticeably to housing prices. Taken together, these results point to a potentially important role for house price bubbles in determining housing investment. We examine the economic significance of the connection between house prices and investment by focusing on events since the mid-1980s. We find noticeable, apparently bubble-induced swings in prices and investment in five of the nine US census regions.

Our results also shed light on the importance of credit availability for house prices and housing investment. Some observers have suggested that increased credit availability may have helped inflate house prices across the OECD during the mid-1980s (for example, Borio et al. (1994)). A separate literature suggests that changes in credit availability can affect investment, owing to informational asymmetries between borrowers and lenders (Hubbard (1996)). The evidence presented here does not point to any link between mortgage credit availability and either house prices or housing investment in the US.

The paper is organized as follows. Section 1 focuses on house prices in the 50 US states, developing evidence that both economic fundamentals and speculative bubbles played important roles over 1973-96. Section 2 focuses on housing investment, and estimates how this component of real activity is influenced by house prices; the section also considers the potential magnitude of bubble-induced swings in investment. Section 3 discusses the policy implications of our results.

1. Speculative bubbles and US house prices?

This section examines whether US house prices were subject to speculative bubbles over 1973-96. After reviewing what is known about bubbles in general, we provide some evidence supporting their existence in the house prices of many US states in the late 1980s. For convenience, we treat bubbles as sustained price rises above fundamentally-determined values, consistent with their common image; however, negative bubbles are certainly also conceivable. The section ends with caveats about the difficulty of verifying the presence of speculative bubbles.

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Speculative bubbles: an overview

Since 1852, when Charles McKay documented some dramatic speculative bubbles in his *Extraordinary Popular Delusions and the Madness of Crowds*, most observers have attributed speculative bubbles to irrational investor behavior. To understand an irrational bubble, it is important to note, first, that prices sometimes rise for perfectly sensible reasons, such as strong economic growth. If such a rise lasts long enough, naive investors may gain confidence that prices will continue to rise. Based on this confidence, they may direct more funds to the market, propelling prices up farther and helping attract more investors. In this way, price rises come to depend on the expectation of further price rises, eroding the line between price levels and fundamentals. Over time, informed investors increasingly realize that prices are unreasonably high and begin pulling funds out. This slows the rise of prices, which in turn, discourages the less informed investors. Eventually, confidence and prices collapse together.

Based on an extensive historical survey, Kindleberger (1978) constructs a more detailed theory of the development of irrational speculative bubbles. Since Kindleberger's book, economists have learned a number of cautionary lessons about speculative bubbles. First, speculative bubbles need not be irrational (Blanchard (1981)). It is possible that speculators are aware of the misalignment between prices and fundamentals, but continue to invest quite rationally on the expectation that the bubble is unlikely to burst. Even so, the fact that irrational speculative bubbles are regularly generated in experimental asset markets (Smith et al. (1988)) does suggèsts that irrationality could be an important factor in real-world asset market bubbles.

Second, there can be extreme price cycles in which prices never depart from their fundamental values. A good example of extreme asset market behavior that might in fact have been consistent with fundamentals is found in the "Tulipmania" of 1634-37, when prices for rare tulip bulbs in the Netherlands skyrocketed and then crashed. Garber (1989) shows that the price behavior of rare bulbs appears consistent with the underlying fundamentals, and that such a precipitous rise and decline was not uncommon for new strains of bulbs. Since extreme price movements can be driven by fundamental factors, it is not possible to prove that a specific historical episode was truly a bubble (Hamilton and Whiteman (1985)). After all, some unrecorded but sensible consideration (an "unobserved fundamental") could have motivated investors at the time; and the absence of any such consideration can never be conclusively established. Nonetheless, it seems difficult to discover what fundamental consideration may have driven some apparent bubbles, such as the 1987 stock market crash (Shiller (1989)).

Speculative bubbles in US house prices: econometric tests

The hypothesis that US house prices have experience speculative bubbles is certainly not new: evidence suggesting the presence of bubbles in regional US housing markets is presented in Poterba (1991) and Abraham and Hendershott (1993, 1994). Muellbauer (1996) presents evidence that house-price bubbles have also been present in the United Kingdom, and Higgins and Osler (1997) for the presence of bubbles in many OECD housing markets during the late 1980s. We develop our own econometric evidence of this point in order to facilitate later analysis of the effects of price bubbles on housing investment.

Our house price variable represents median new house prices by state for 1973-96. We combine this with other state-level variables to form a panel of annual data for the 50 states covering 1973-96. We divide our independent variables into fundamental and non-fundamental house price determinants. Since we cannot directly measures the presence of speculative bubbles, our evidence concerning the importance of speculative bubbles is necessarily indirect in nature.

Fundamentals: One simple and robust model asserts that an asset's price should equal the present discount value of the associated income stream:[1]

$$House\ \ price_t = rent_t^e + \frac{rent_{t+1}^e}{(1+r_t)} + \frac{rent_{t+2}^e}{(1+r_t)^2} + \ldots$$

Here, *e* indicates that the share price is based on the *expected* value of future rents, and *r* represents an appropriate discount rate. We derive an estimating equation consistent with this theory by restating it as follows:

$$House\ \ price_t = rent_{t+1}^e + \frac{House\ \ price_{t+1}^e}{(1+r_t)}$$

The expression states that the current price of a house should equal expected rents over the coming year plus its own value one year hence discounted to the present.

Table 1

Panel unit root tests

Panel unit root tests were implemented for a model containing state-specific dummy variables and rely on the critical values reported by Levin and Lin (1992). The null hypothesis is that the variable in question is I(1). For the variables included in our empirical analysis, we report below whether the unit root null is rejected and, if so, at the 1, 5, or 10% level. Level variables are measured in logs, except for the real mortgage interest rate and the user cost of housing.

Level variables	**Significance level**
Affordability	1%
Authorizations	5%
Consumer price index	not rejected
Construction costs (real)	not rejected
Employment	not rejected
House prices (real)	1%
Mortgage interest rate (real)	1%
Mortgage originations (real)	10%
Per capita income (real)	not rejected
User cost of housing	1%

Differenced variables	**Significance level**
Affordability	1%
Authorizations	1%
Consumer price index	1%
Construction costs (real)	1%
Employment	1%
House prices (real)	1%
Mortgate interest rate (real)	1%
Mortgage originations (real)	1%
Per capita income (real)	1%
User cost of housing	1%

[1] See, for example, Copeland and Weston (1988), pp. 20-2, or Brealey and Myers (1987), pp. 44-5.

The formula suggests that our empirical model should include expected house rents, current mortgage interest costs, and expected future house prices. State-level data on actual house prices are readily available, and national information on the cost of mortgages can be adjusted to state-specific real values using state CPIs. However, data concerning expected rents and expected future house prices are not available. We estimate the influence of expected rents implicitly, using three factors likely to determine rents: per capita income, employment, and construction costs. Following convention, we approximate expected house prices as an autoregressive process; experiments indicate that two lags are relevant.[2]

To accommodate the stationarity properties of these data (described in Table 1), we take the growth rate of real house prices, rather than their level, as the dependent variable. Accordingly, the fundamental house-price determinants included in our panel are growth rates of the variables listed above. (The Appendix describes our data sources, and provides further details concerning variable measurement.) We include state dummies to capture persistent unmodeled or idiosyncratic factors that might vary by state. Aside from the dummy-variable coefficients (which amount to state-specific intercepts), the estimated coefficients are assumed to be the same for each state. Limiting our analysis to the fundamental determinants of house prices would lead to a regression specification such as the following:

$$\Delta HP_{it} = \alpha \Delta YD_{it} + \beta \Delta EM_{it} + \gamma \Delta CC_{it} + \mu \Delta r_{it} + \nu \Delta Hp^{e}_{it+1} + s_i + \varepsilon_{it}$$

where i indexes states and t time, while HP represents house prices, Y per capita disposable personal income, EM employment, CC construction costs, r real mortgage interest costs, s state-specific factors (constant over time), and ε is a residual.

Non-fundamentals: One of our central theses is that house prices are affected by speculative bubbles. The total contribution of non-fundamental forces which includes the contribution of speculative bubbles – could be assessed by examining the residuals from the above regression. However, since we are primarily interested in the contribution of speculative bubbles, our strategy is to estimate the influence of all potential non-fundamental forces individually, which implies regressions of the following form:

$$\Delta HP_{it} = \alpha \Delta YD_{it} + \beta \Delta EM_{it} + \gamma \Delta CC_{it} + \mu \Delta r_{it} + \upsilon \Delta HP^{e}_{it+1} + \sum_{j=1}^{k} \Psi_j NF_{ijt} + s_i + \varepsilon_{it}$$

where NF_j represents any non-fundamental force.

In addition to speculative bubbles, we focus on two other non-fundamental factors that may have influenced house prices in some states: credit availability and overbuilding. We discuss our measures of these two additional non-fundamental forces before returning to consider speculative bubbles.

A role for **credit availability** in determining house prices is suggested in Borio et al. (1994), who argue that the rapid rise of house prices around the OECD during the late 1980s was due in part to rapid contemporaneous growth in mortgage credit. Corresponding to the possibility that credit growth fueled the asset price spikes, the later price declines could be attributed to a "credit crunch" in the early 1990s. Such a credit crunch, if it occurred, might, among other factors, owe to BIS bank capital standards, imposed beginning in 1988 (Bernanke and Lown (1992)). The idea that credit dynamics could affect house price growth is closely related to the bubble hypothesis: a bubble occurs whenever asset prices experience a sustained rise beyond the levels justified by fundamentals, and this remains true even if the bubble is accompanied or fueled by rapid credit growth.

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[2] Note that there is an inherent difficulty in modeling expectations. We are not claiming that expectations are formed rationally. On the other hand, we model them as though they were formed rationally given lagged price information. This difficulty would not arise, of course, if survey data on house price expectations were available.

To assess the contribution of mortgage credit availability to house price growth, we would ideally include a measure of the growth in mortgage credit outstanding, by state, for our entire sample period. The available state-level data fall short of this ideal in two ways, however. First, they are only available from 1983 through 1993. Second, they cover mortgage originations, which include refinanced mortgages as well as new ones. To deal with this second problem, we measure state-level originations as deviations from the national average.

According to the overbuilding hypothesis, excessive investment during the 1980s could have left a substantial backlog of unoccupied new homes in some areas. This role for overbuilding in deflating asset prices is also compatible with the bubble hypothesis: just as excessive optimism leads investors to raise asset prices past the levels justified by fundamentals, builders might construct homes beyond levels justified by a sober analysis of potential demand growth. Overbuilding is considered "non-fundamental" because, in an efficient market, prices would adjust swiftly to supply, and any lagged supply variable would be uncorrelated with current price changes. Because our data are measured annually, the speed at which prices would be required to adjust to meet this criterion would not be great. In the absence of a natural measure of overbuilding, we experiment with two different proxies for it. The first is the ratio of cumulative housing authorizations to a state-specific trend. The second is the ratio of housing stock to population, where the housing stock is estimated using the perpetual inventory method, since state-level housing stock data apparently do not exist.

Unfortunately, there is no true "measure" of the forces behind a speculative bubble. However, there are two properties of speculative bubbles that we can use to evaluate whether they might have existed. First, the farther prices rise relative to fundamentals as a bubble takes hold, the farther they must fall relative to fundamentals later on. Second, during a bubble, the initial rise of prices above fundamentals, as well as the subsequent decline, should be fairly monotonic. We attempt to capture the first property directly, and the second property by examining regression residuals.

The first property of bubbles implies that, on average, a positive gap between prices and their fundamentally-determined values should be associated with subsequent price declines. Further, the larger the gap, the larger the later decline. One way to capture this property would be to include the lagged level of real house prices as an explanatory variable. Since our regression includes state dummies, this would in effect measure house prices as deviations from state-specific averages. If house prices are characterized by constant state-specific fundamental values, any speculative bubble component of prices should be captured by a negative coefficient on lagged prices. There is some empirical support for this crude view of house price determination. In particular, standard panel unit root tests indicate that state real house prices are I(0), narrowly rejecting the null hypothesis of a unit root at the one percent level (Levin and Lin (1992, 1993)). Hence, departures in real house prices from their state-specific averages tend to erode over time, suggesting the presence of constant state-specific fundamental values.

Even so, we are not fully convinced that the lagged house price level represents an appropriate measure of this first property of speculative bubbles. An important source of our skepticism is the fact that real house prices display a clear upward trend, rising by 26.7% at the national level from 1973 to 1996. Moreover, the literature on testing for unit roots in a panel setting remains in flux, with standard tests recently criticized for rejecting the unit-root null too frequently (O'Connell (1997)).

As an alternative, we use the lagged ratio of real house prices to real disposable income. This variable is also used by Muellbauer (1996), who labels it "affordability." Since the variable is lagged by a full year, it would not affect price growth in a fully efficient market. Further, its inclusion has a natural economic interpretation consistent with the presence of speculative bubbles: if houses become too unreasonably expensive, demand will dry up, forcing prices back down again. This variable is more unambiguously stationary than real house price levels: standard panel unit root tests reject the I(0) null at better than the 0.1% level. Moreover, the variable displays no trend, remaining virtually unchanged at the national level since 1973.

Relying on affordability to capture this first property of speculative bubbles is akin to treating house prices and disposable income as cointegrated. Affordability could then be seen as a quasi-error-correction term in the regression for house price growth. Though appealing, this interpretation of our regression equation is not econometrically reliable, for two reasons. First, affordability could not be an exact error correction term in our specification, since the cointegrating relationship implied by that specification includes several variables, not just per capita income. Second, little is known about estimating cointegration-ECM relationships in a panel setting. The literature on testing for cointegration in a panel setting is in its infancy and no clear consensus has emerged regarding appropriate test techniques or significance levels (Pedroni (1995, 1997)). Beyond this, there is apparently no work which estimates a panel error-correction model.

Table 2

Panel regressions of annual house price growth by US state, 1973-96

	Regression 1	**Regression 2***	**Regression 3**	**Regression 4**
Fundamentals				
Per capita income growth	0.26 (2.62)	0.44 (3.17)	0.26 (2.72)	0.26 (2.65)
Employment growth	0.30 (2.98)	0.34 (2.80)	0.30 (2.97)	0.31 (3.04)
Growth in construction costs	0.76 (8.43)	0.27 (2.18)	0.77 (8.24)	0.77 (8.36)
Real mortgage interest growth	-1.04 (-9.07)	-0.90 (-3.68)	-1.05 (-9.11)	-1.04 (-9.07)
Expected house price appreciation	0.43 (4.41)	0.72 (3.70)	0.40 (3.46)	0.43 (4.46)
Non-fundamentals				
House "affordability", lagged (Ratio of price to per capita income)	-16.30 (-6.89)	-30.36 (-5.47)	-16.77 (-6.54)	-16.35 (-6.87)
Growth in mortgage originations, lagged		-0.00 (-0.20)		
Overbuilding 1 (Deviation of cumulative housing authorizations, from state trend)			0.40 (0.44)	
Overbuilding 2 (Housing stock/population ratio)				2.62 (1.22)
Number observations	1,071	561	1,020	1,020
R-bar squared	0.31	0.34	0.31	0.31
Durbin-Watson statistic	2.02	1.99	2.02	2.02

Note: t-statistics in parentheses.

* The sample size is smaller for this regression because mortgage origination data only span 1983-93.

Although affordability may help us capture the first property of speculative bubbles, it will not fully capture the second property of speculative bubbles listed above, that bubble-induced price movements should include a fairly monotonic rise above fundamentals followed by a fairly monotonic decline back towards fundamentals. We will use an analysis of the regression residuals to capture this second property. There are other aspects of speculative bubbles that affordability may not capture at all. For example, bubbles may be based on irrational expectations of continuously rising prices, but there is no way to capture that irrationality in the absence of survey data on expectations.

Results: The results of our analysis are presented in Table 2, where we show a few different versions of our baseline regression. All the regressions display reasonably high explanatory power with low residual autocorrelation. The estimated coefficients for the fundamental variables all have the expected signs and have economically sensible magnitudes, and they are all statistically significant.

With regard to possible non-fundamental influences on house prices, the only non-fundamental variable with any apparent explanatory power is "affordability," which we interpret as capturing the fact that prices inflated by bubbles must eventually return to fundamental values. The coefficient on affordability is consistently negative as expected, and significant. Its magnitude, which varies only slightly across regressions, implies that a 10 percentage point deviation of the house-price/per capita-income ratio from its state-specific average is typically followed by a 2 percentage point decline in house prices the following year.

The coefficients on both mortgage credit availability and overbuilding are statistically insignificant and have unexpected signs. The statistical significance of the fundamental variables declines when mortgage credit is added to the model, but unreported results indicate that this is largely due to the constrained sample size. The exclusion of mortgage credit and overbuilding has little effect on the coefficients of the remaining variables or the other properties of the regressions. Regression 4, which is our preferred specification, includes only fundamentals and affordability.

Speculative bubbles in US house prices: 1982-93

The results so far support evidence from other studies that speculative bubbles could affect US housing markets. We have not yet examined, however, the second property of speculative bubbles listed above: that prices will tend to rise monotonically and then fall monotonically relative to fundamentals. Nor do the results tell us whether speculative bubbles have been important in economic terms. To address these issues, we now focus on 1984-93, and ask whether house prices in some regions overshot fundamental values for extended periods, and subsequently suffered sustained declines.

A quick review of the aggregate data suggests that US house prices movements were generally quite moderate during 1984-93: aggregate (population-weighted) real house prices rose 26% during 1982 to 1989, a period of rapid GDP growth (GDP itself grew over 30%), and fell just 4% during the slow-growth period from 1989 to 1993 (see Table 3). However, these moderate aggregate price movements mask dramatic regional swings. In New England, for example, real house prices rose 49% during 1982-89, and fell 17% during 1989-93. Large price movements were also observed in the Mid-Atlantic, Mountain and Pacific census regions.[3] For all four of these regions, price rises exceeded the national average over 1982-89 and price declines exceeded the national average over 1989-93. Since speculative bubbles tend to be identified with periods of extreme price movements, one might venture a preliminary guess that these four regions experienced such bubbles during 1984-1993.

The total amount of house price growth not determined by fundamentals can be estimated on the basis of Regression 4 of Table 2 as the regression residuals plus the contributions of deviations of affordability from its state-specific average. The cumulated value of these non-fundamental annual price movements, which are shown in Charts 1A through 1D for the nine US census regions, represents our measure of the total departure of prices from fundamental values.

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3 Our regional definitions are taken from the US Census Bureau.

Table 3

Cumulative growth in real house prices and house authorizations, by region

	House prices		House authorizations	
	1982-89	**1989-93**	**1982-89**	**1989-93**
US	26.2	-4.2	41.4	10.9
New England	49.0	-17.1	56.2	-16.4
Mid-Atlantic	40.3	-12.7	80.7	-21.3
South Atlantic	19.2	1.9	46.0	6.0
E. N. Central	21.6	-1.7	113.4	21.2
E. S. Central	19.8	5.2	55.2	35.1
W. N. Central	18.9	1.8	41.4	30.1
W. S. Central	13.5	0.4	-89.9	53.1
Mountain	38.0	-7.6	-18.8	74.4
Pacific	27.7	-9.4	91.4	-43.3

Notes: Prices correspond to state median house price data deflated by state CPIs. Authorizations correspond to new single-family homes.

The five regions included in Chart 1A and 1B are those where the non-fundamental component of house prices is consistent with the second property of bubbles listed above: the component rises consistently and substantially during the late 1980s, and then falls consistently during the early 1990s. These five regions include the four mentioned above, in which average price changes were more extreme than the national average over 1984-93, plus the "East-South Central" region, which includes Louisiana and Texas, among other states. The modest size of the apparent speculative bubble, which peaked well before the national house price peak in 1989, suggests that most of the price dynamics in this region were driven by fundamental forces such as swings in oil prices. The remaining four regions are shown in Charts 1C and 1D, where it can be seen that the influence of non-fundamental forces was consistently small and did not conform to the up-down bubble profile.

In our introduction, we suggested that policymakers may be concerned about the "hangovers" associated with asset market bubbles. One such hangover would be the price declines associated with the collapse of such a bubble. These reduce homeowner wealth and they can also lead to increased defaults, if some homeowners find that their mortgages exceed the value of their house. The price declines can even impede the proper functioning of labor markets, to the extent that homeowners feel trapped in their existing home and unable to take advantage of new job opportunities elsewhere. Our results allow us to estimate crudely the extent to which the early 1990s' house price deflation is attributable to speculative excesses in the late 1980s.

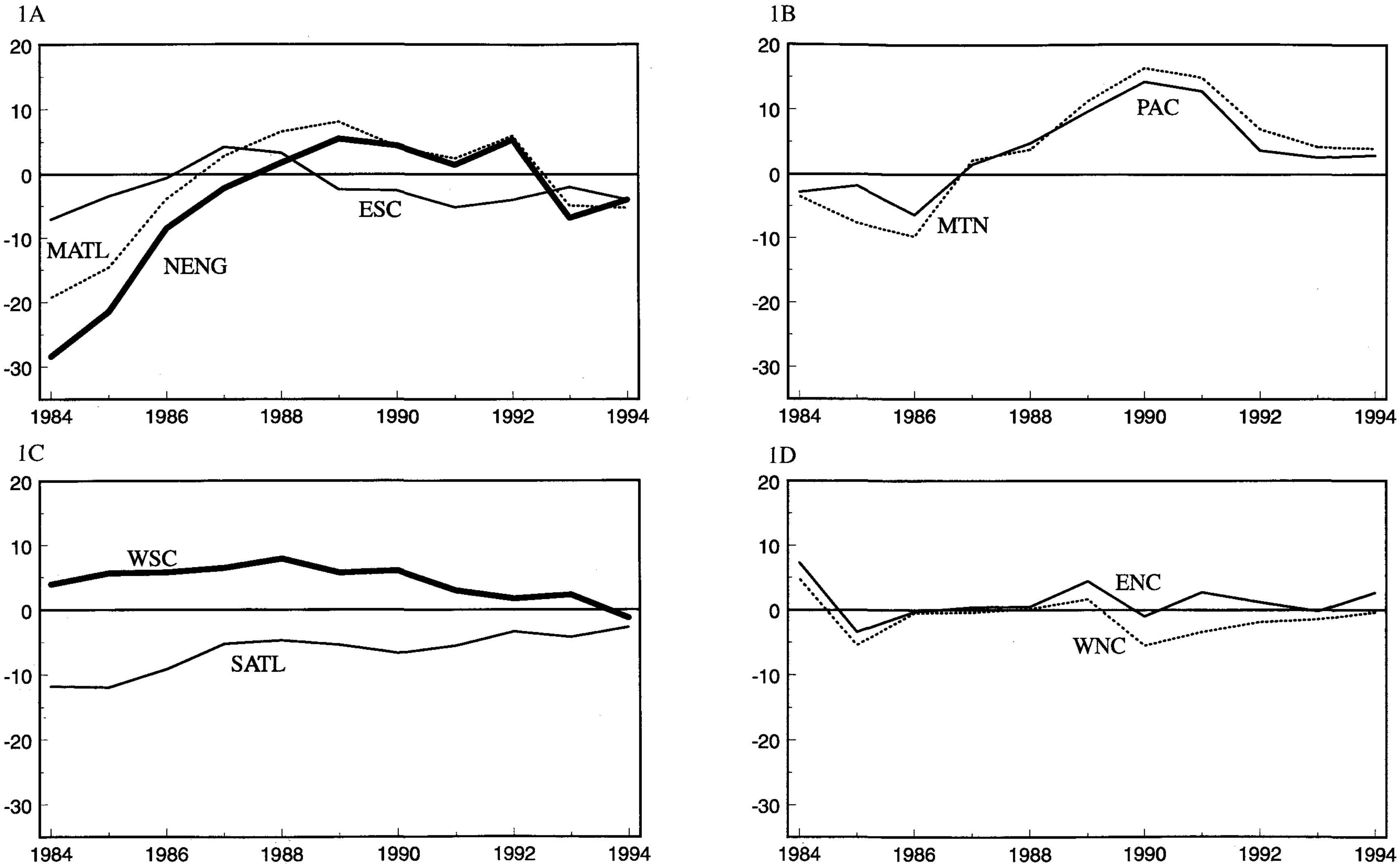
Table 4 shows total house price declines over the four years following regional peaks, the amount of that decline attributable to affordability, and the amount attributable to non-fundamental forces more generally (affordability plus residuals). The measures associated with affordability correspond to the first property of speculative bubbles listed above: the fact that, the further prices initially rise, the further they ultimately must fall.

In the five regions where bubbles were apparently important, real house prices declined by almost 10% over the four years following their regional peaks. Affordability itself accounts for an average decline of 6.4% in these five regions. In the other four regions, where house prices declined only 2.1% on average in their four post-peak years, affordability accounts for virtually none of the price declines.

The total effect of non-fundamental forces, meanwhile, was to depress prices by almost 12% in the regions identified as most likely to have experienced bubbles, about 2 percentage points

Chart 1

Non-fundamentally determined house price growth, in percentages



NENG: New England; MATL: Mid-Atlantic; ESC: East South Central; MTN: Mountain; PAC: Pacific; SATL: South Atlantic; ENC: East North Central; WNC: West North Central; WSC: West South Central.

Table 4

Real house price declines, four years from peak

Price declines	**Bubbly ?**					
	New England	Mid-Atlantic	Pacific	Mountain	E. S. Central	Average
Actual	-17.1	-12.7	-7.3	-6.9	-4.2	-9.7
Due to affordability	-7.9	-7.4	-5.9	-7.5	-3.1	-6.4
Due to non-fundamental factors	-12.3	-12.9	-11.6	-12.1	-9.8	-11.9
	Non-bubbly ?					
	E. N. Central	South Atlantic	W. N. Central	W. S. Central	Average	
Actual	-1.7	-1.7	1.8	-5.0	-2.1	
Due to affordability	-0.4	1.2	0.5	-1.7	-0.0	
Due to non-fundamental factors	-5.1	1.0	-3.6	-6.2	-3.1	

Notes: Prices represent median state house prices deflated by state CPIs. Estimated contributions of "affordability" and "non-fundamental factors" are based on regression 4 from Table 2.

more than the actual price decline.[4] Similarly, non-fundamental forces reduced house prices by about three percent in the other four regions, about one percentage point more than the actual decline. Although it is difficult to know how seriously to take differences of this magnitude, they do suggest that fundamental forces tended to support prices in these regions over these four years, and that price declines might have been more extreme had only the estimated non-fundamental forces been at work.

So far we have provided graphical and statistical evidence indicating that US house prices were susceptible to speculative bubbles over 1973-96. Further, we showed that speculative price bubbles may have arisen in the north-east, the far west, and the "East-South-Central" regions during the late 1980s, and that those bubbles may have been an important sources of local house price weakness in the early 1990s. The results do not *prove* that bubbles were important: as mentioned earlier, it is impossible to know for certain whether a given asset boom truly represents a speculative bubble, as some unobserved fundamental could always be driving prices. Nevertheless, the results do place the alternative, non-bubble hypothesis in sharper relief, by limiting the unobserved or unmeasured fundamentals consistent with observed house price behavior. In particular, if speculative bubbles do not explain the boom/bust cycles beginning in the mid-1980s, the unobserved or unmeasured fundamentals which do explain the cycles must have deteriorated most sharply in precisely those regions where they previously improved most sharply. For example, if state income taxes were a candidate unmeasured fundamental, this would require that such taxes rise the most in the early 1990s in those states where they declined the most in the late 1980s. The alternative, non-bubble hypothesis thus appears to require an unlikely confluence of events.

4 Qualitatively similar results are obtained if we look at price changes over the two years following the peak.

2. House prices and housing investment

In this section we turn our attention from house prices to house investment, and ask whether house price bubbles might be an important determinant of real activity. The section begins with a general discussion of the connections between house prices and housing investment, none of which imply irrationality among home builders. We then estimate this relationship empirically for the 50 US states. Finally, we evaluate the extent to which growth in housing investment may have been depressed in the early 1990s amid the hangover from earlier speculative excesses in the housing market.

House prices and housing investment: an overview

There are several possible connections between house prices and housing investment. First, expected house price appreciation affects the attractiveness of housing as an investment asset, with potential builders responding to the prospect of capital gains or losses. Second, the *level* of house prices might also discourage housing investment, even if potential investors do not expect house prices to change. Our focus in this paper is on the second of these connections.

House price levels can affect construction directly and through their effect on credit availability. The direct effect on construction works through the mechanism identified in Tobin's q theory of investment (1969): potential builders are unlikely to engage in speculative construction, and prospective homeowners will prefer to buy an existing home, if house prices are lower than the cost of construction.

The indirect effect of house prices, which works through credit availability, can affect housing investment in multiple ways. Declining house prices lower homeowners' net worth, and some homeowner will not have sufficient assets for a down payment on another house, if they are inclined to move or to trade up. Other homeowners may find themselves saddled with mortgage obligations greater than the value of their home, perhaps inducing default. Increased defaults reduce lenders' capital, possibly reducing the supply of mortgage credit.[5]

House prices and housing investment: empirical tests

To evaluate the strength of the connection between housing investment and house prices we develop an empirical model based primarily on the neoclassical model of business fixed investment, by Jorgenson (1971) and others. As modified to apply to housing investment, neoclassical theory suggests that investment in state i in year t should be positively related to expected future rents and the existing stock of housing (via depreciation), and negatively related to the user cost of capital. More recent theories which relate investment to asset prices via credit markets, described above, suggest that housing investment should be positively related to the level of house prices and to mortgage credit availability. Casual empiricism suggests that, in some regions, overbuilding may have been a determinant of housing investment during our period of interest.

We use authorizations for the construction of new single-family houses as our measure of housing investment, a choice dictated by data availability.[6] As earlier, we allow for the influence of

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5 This discussion condenses and, inevitably, simplifies an enormous literature on the subject of asymmetric information and the role of credit in business cycles. Surveys can be found in Hubbard (1996), Bernanke et al. (1996), Kashyap and Stein (1996), Bernanke (1993), and Gertler (1988).

6 Our measure of housing authorizations differs somewhat from "residential construction," the measure of housing investment included in the national income and product accounts. First, our measure does not include any replacement investment. Second, our measure excludes multifamily homes, condominiums and apartments. We chose to focus on

expected rents implicitly, by including factors that should determine them, specifically per capita income and employment. We again use mortgage originations to proxy for mortgage credit availability. In the present context, the influence of overbuilding should be captured by a negative coefficient on the lagged capital stock.

The user cost of housing is determined by mortgage interest rates, expected depreciation, expected capital gains and the cost of construction.[7]

$UC_t = CC_t\,[r_t - depreciation^e{}_t - capital\ gains^e{}_t]$

where UC represents the user cost of capital. For construction costs (CC), we use a national measure deflated by state CPIs. State mortgage interest rates are estimated as the national mortgage rate minus state CPI inflation over the past year. Depreciation is taken to be 3.5% per year, following Summers and Heston (1995 a, b). Expected capital gains are represented by the estimate of expected house price appreciation discussed in the previous section.

In constructing our estimating equation, we take annual *growth* in (log) housing authorizations as our dependent variable; the independent variables described above thus also appear as changes or growth rates. We approximate the change in state housing stocks with authorizations themselves. State- and time-specific effects are also included, as is a lagged dependent variable, intended to capture the influence of unmodeled forces. We use changes in actual mortgage originations, rather than their deviations from national averages, because the effect of national refinancing trends should be captured by time dummies.

Our choice of functional form is informed by the fact that panel estimates which include a lagged dependent variable along with state-specific effects are biased, especially when the time dimension of the panel is small or moderate. Unbiased estimators have been developed by Anderson and Hsiao (1981) and Arellano (1989), who apply IV methods to differenced variables, using appropriate lags as instruments.[8] Our estimating equation is then given by:

$$\Delta HA_{it} = \alpha\Delta HA_{it-1} + \sum_{j=0}^{n}\beta_j \Delta YP_{it+j} + \sum_{j=0}^{n}\gamma_j \Delta EM_{it+j} + \sum_{j=0}^{n}\mu_j \Delta UC_{it-j} + \sum_{j=0}^{n}\eta_j MO_{it-j} + \sum_{j=0}^{n}\nu_j HA_{it-1} + s_i + \lambda_t + \varepsilon$$

Here, HA represents housing authorizations, MO mortgage originations, s a state dummy (constant across time), λ a time dummy (constant across states), and ε is the residual.

The regression results are presented in Table 5. Per capita income growth is excluded in all results since this variable was consistently statistically insignificant. We found that one lag of all variables was sufficient, which is not surprising since house construction generally takes less than a year to execute. Mortgage credit is excluded in the first column since this variable is available for only about half of our sample period. The second regression suggests that mortgage credit availability is not an important determinant of housing authorizations, once other fundamental factors are accounted for.

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single-family units to preserve compatibility with our measure of house prices. It is possible, of course, that our data include projects which were authorized but never carried out.

7 Here, again, we allowed the effect of mortgage interest rates and expected house price appreciation to enter the regression separately, but the coefficients were extremely close and statistical tests indicated that they should be combined.

8 Consistent estimation requires that the dependent variable be lagged twice before inclusion as an instrument. Differencing the original investment equation, as we do in moving from the expression for HA_{it} to that for ΔHA_{it}, produces a moving-average error term correlated with HA_{it-1}. For this reason, HA_{it-2} was the most recent lag of investment used as an instrument, and we used techniques described in Newey and West (1987) to control for the moving average component of the error term.

Table 5

Panel regression of annual growth in house authorizations by US State, 1973-96

	Regression 1	Regression 2*
Lagged growth in real house prices	0.34 (3.21)	0.24 (1.58)
Change in employment	1.54 (2.25)	1.08 (2.98)
Change in user cost of housing capital	-1.86 (-1.77)	-1.55 (-1.23)
Lagged authorizations	-17.32 (-4.78)	-18.49 (-4.74)
Lagged dependent variable	0.55 (5.09)	0.63 (19.24)
Lagged growth in mortgage originations (credit availability)		0.03 (0.99)
Number of observations	1,019	561
R-bar squared	0.57	0.30

Notes: Authorizations and prices correspond to single-family homes. The user cost of housing capital incorporates construction costs, mortgage interest rates, depreciation and expected house-price appreciation. t-statistics in parentheses.

* The sample size is smaller for this regression because mortgage origination data only span 1983-93.

In consequence, we concentrate on the first regression, which generally supports the theoretical predictions discussed above. Coefficients on all fundamental investment determinants have the expected sign and sensible magnitudes. All variables are significant at the 5% level, except the user cost of housing which has a marginal significance level of 8%. The significantly negative coefficient on lagged authorizations suggests a potentially important role for "overbuilding."

For our purposes, the relationship of greatest interest is between house prices and investment. The regression results point to a reasonably strong link between the two variables, with a point-in-time elasticity of 0.34 and a long-run elasticity of 0.77.[9] If our finding that mortgage origination growth is not economically important for housing investment is correct, one can infer that house prices primarily influence housing investment through the direct effect (analogous Tobin's q theory), rather than through their influence on credit.

House prices and housing investment

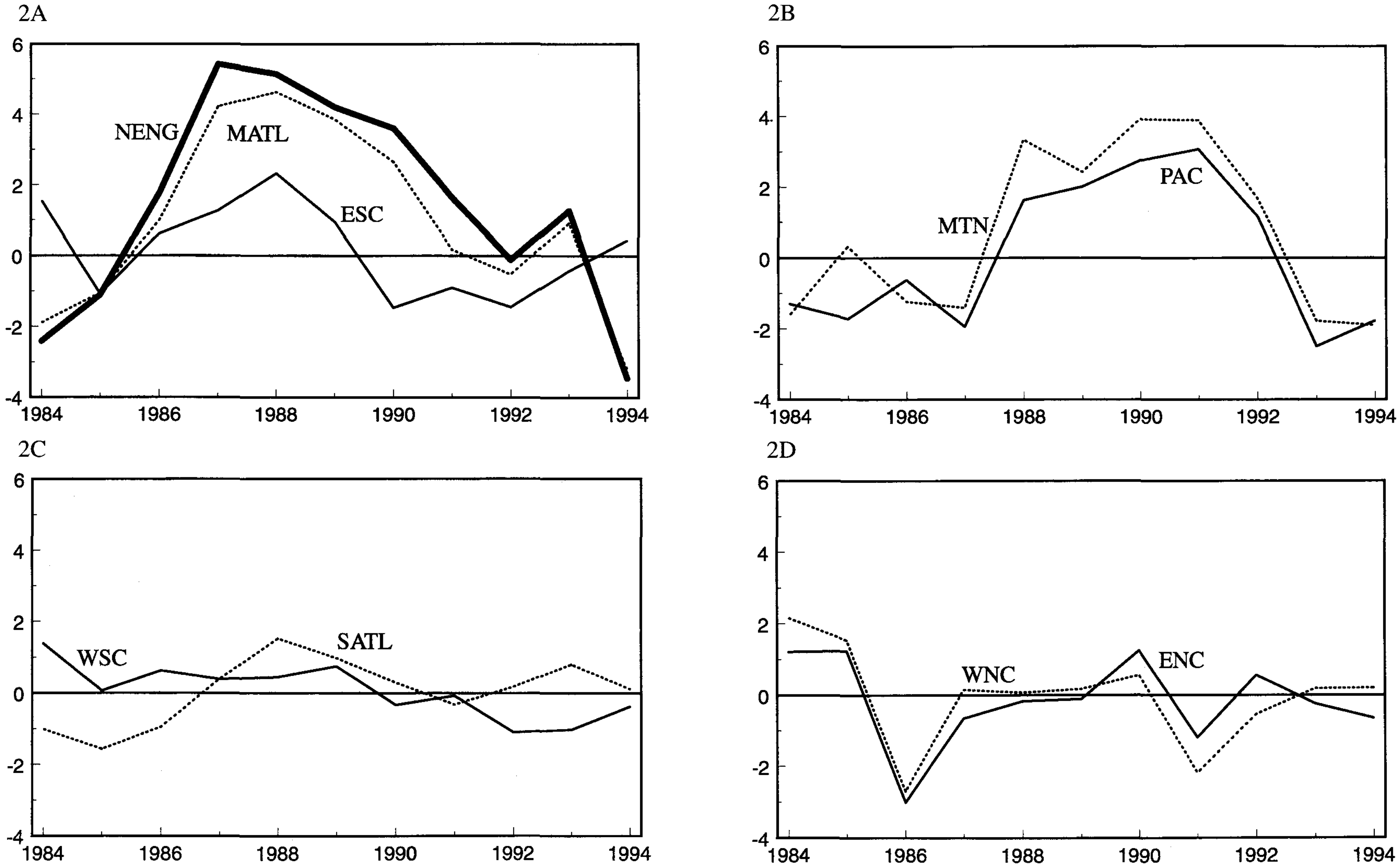
To examine whether house prices have indeed been important determinants of housing investment, we focus once again on the period from the mid-1980s to the mid-1990s. At the national level, housing investment growth peaked in 1986 at 9%.[10] Following that peak, housing investment growth turned negative, and remained so through 1991. House prices, however, would only have been a drag on investment growth following their own peak in 1989. Our estimates suggest that, amid

[9] The short and long-run effects of house price growth differ because price growth also influences current investment indirectly through lagged investment.

[10] Note that the peak of housing investment precedes the peak in house prices by a few years. This is typical, and it highlights the important fact that house prices are just one determinant of housing investment.

Chart 2

Housing authorization growth attributable to non-fundamentally determined house price growth, in percentages



NENG: New England; MATL: Mid-Atlantic; ESC: East South Central; MTN: Mountain; PAC: Pacific; SATL: South Atlantic; ENC: East North Central; WNC: West North Central; WSC: West South Central.

the recession of 1990-91, house price movements reduced housing investment growth by 0.8 percentage points in 1990 and by a further 1.8 percentage points in 1991. By 1992, when the economy confronted mysterious "headwinds" as it tried to recover, house price movements may have reduced housing investment growth relative to its 1989 level by a full 3.9 percentage points. In short, the estimates suggest that house prices declines may have noticeably reduced housing investment growth during the early 1990s.

Our central thesis is that speculative bubbles in house prices can affect housing investment. To evaluate this thesis, we combine the measure of non-fundamental house price movements developed in Section 1 with the regression estimates of the effect of house prices on housing investment. This allows us to calculate the contribution of non-fundamental price movements to housing investment growth.

Charts 2A through 2D show the estimated influence of non-fundamental house price movements on housing investment over 1983-93, broken down by region. In the regions identified previously as possibly experiencing speculative bubbles, these movements boosted housing investment on the price upswing and depressed housing investment on the downswing. This "hangover" effect on housing investment was apparently quite substantial. For example, during the five years following the regional house price peak in 1989, cumulative housing investment growth in New England was slowed by more than seven percentage points. In the Mid-Atlantic states, the corresponding figure is 5.9%. Across these five regions, investment growth was reduced by 5.0 percentage points on average due to non-fundamental price movements following price peaks. For the remaining four regions, the corresponding figure is 1.1 percentage points.

At the national level, diversity across the US states mutes the effect of non-fundamental forces on housing investment. The estimates suggest that non-fundamental price movements reduced cumulative growth in national housing investment by 2.7 percentage points over the five years following the national house price peak in 1989. Of this, a full 1.1 percentage point took place in the recession year of 1991 and an additional 1/4% the following year. While these effects are moderate in scale, they do suggest that a slowdown in housing construction associated with non-fundamental house price movements could have contributed to the early 1990s recession, and to the "headwinds" that slowed the ensuing recovery.

Conclusion

Our paper presents evidence that speculative bubbles in US house prices can effect housing investment. Based on annual data covering the 50 US states, we derive evidence from two separate panel regressions suggesting, first, that non-fundamental forces have had a significant influence on house prices; and second, that these in turn have had a significant influence on investment. Taking these results together, non-fundamental movements in house prices appear to have had a noticeable impact on housing investment. We use the econometric results to show that the tumbling house prices and anemic investment observed in many regions during the early 1990s could have represented, in part, the "hangover" from speculative house price bubbles in the late 1980s.

The idea that asset market behavior could have substantial effects on real economic activity is not new: as early as 1933, Irving Fisher claimed that debt deflation contributed importantly to the great depression. More recently, economists have fleshed out our theoretical understanding of these real-financial linkages, and much evidence has accumulated suggesting the importance of such linkages in earlier historical episodes.[11] Our results support the idea that asset price developments continue to affect real activity.

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[11] For recent reviews on this topic, see Bernanke and Gertler (1995) or Bernanke et al. (1996). For additional empirical evidence, see Hubbard (1994).

Beyond this, our evidence suggests that asset price movements not based on fundamentals can have important implications for economic stability, which raises an important policy question: Should governments try to contain or to prevent speculative asset price bubbles? We introduce this issue here without taking a stand on its resolution. In considering this question, governments could choose among policy alternatives including monetary policy, tax policies, or regulation.

Monetary policy could be tightened in response to excessive speculative activity: higher interest rates should directly reduce equilibrium asset prices. Further, the associated decline in the value of assets used as collateral would discourage the heavy borrowing typically associated with speculation. Though this policy is fairly certain to have the desired effect on asset prices if pursued with sufficient vigor, it has the fundamental problem that, if the bubbles are regionally concentrated, as they seem to have been in US housing markets, monetary policies intended to deflate bubbles in some regions would also affect the other regions.

A monetary attack on speculative bubbles would have other problems, as well. Identifying when to intervene would be difficult: for example, though it is by now widely accepted that Japan's stock and property markets were inflated by speculative bubbles in the late 1980s, there was no such agreement at the time. In fact, our best statistical methodologies even have difficulty identifying bubbles in *past* episodes. (One possible solution to this difficulty would be to focus on rapid asset price rises only when they are accompanied by rapid credit growth, as suggested in Schinasi and Hargraves (1993).) Finally, adding speculative asset price movements to the list of intermediate targets for monetary policy could make policy shifts less transparent to the public. One alternative would be for monetary authorities to alert markets to the possibility that asset prices exceed their fundamental values, without actually changing interest rates, a practice commonly referred to as "jawboning."

Tax policies or regulation could attack speculative bubbles in a manner more carefully targeted across the type of market – that is, tax policies could be focused on the housing market or the stock market. However, if applied at the federal level such policies are not likely to be any better targeted regionally than monetary policy. As an example of tax policies, note that capital gains taxes in some countries already attempt to discourage speculative turnover by promoting long-term ownership of investment assets. Requiring hefty minimum down payments on mortgages could also discourage speculative activity. Other regulations could actually prohibit speculative activity, as in some countries where banks have historically been barred from financing commercial building construction until future occupancy is fully committed. Tax policies and regulation could be applied permanently or only when the danger from bubbles appears imminent, much as the Japanese government limited banks' real estate lending during 1990. This, of course, brings back into focus the difficulty of identifying bubbles as they arise.

In short, in deciding whether to attempt to contain or to prevent bubbles, a government must first decide whether there is sufficient information on which to base any policy change. If intervention appears appropriate, it must choose whether the policies should be implemented by the monetary, tax, or regulatory authorities; it must choose the level of government authority most appropriate, and it must choose between permanent measures and those adopted as speculative pressures appear to build.

Our results have direct implications for the connection between monetary policy and asset market bubbles. Some observers have suggested that, as a general principle, easy monetary policy can be an important force behind excessive asset price inflation (Allen and Gale (1997) and Grant (1991)). Others have specified that easy money was in fact an important force behind the asset market booms of the late 1980s (Hoffmaister and Schinasi (1994) and Schinasi (1994)). Monetary policy was, of course, the same for all 50 US states in our panel. Since we find that speculative bubbles were strong in only about half of US states, one might infer that easy monetary policy is not a sufficient condition, and may not even be a necessary condition, for the development of price bubbles.

Data Appendix

Construction costs at the national level are reported in the *Engineering News-Record*, published by McGraw-Hill, Inc., and were taken from the DRI data base. State-level measures of real construction costs were derived by dividing this variable by state-level CPIs.

Consumer price indexes for the 50 US states are reported by the US Department of Labor, Bureau of Labor Statistics, and were taken from the DRI data base.

Disposable income data for the 50 US states are reported the *Survey of Current Business*, published by the US Department of Commerce, and were taken from the DRI data base. Total disposable income was divided by state population to derive per capita disposable income; this in turn was divided by state-level CPIs to derive real per capita disposable income.

Employment data for the US states is reported in *Employment and Earnings*, published by the US Department of Labor, Bureau of Labor Statistics, and were taken from the DRI data base. The data pertain to non-agricultural employment.

Housing authorizations data for the 50 US states are reported by the US Department of Commerce, Bureau of the Census, and were taken from the DRI data base. The data refer to the number of construction permits issued for new single-family homes.

House prices refer to DRI calculations of the median price of new, single-family homes based on national, regional and state-level information on median and mean house prices.

Mortgage interest rates at the national level are reported by the Federal Housing Finance Board, and were taken from the DRI data base. We subtracted state-level CPI inflation during the current year to derive state-level measures of real mortgage interest rates.

Mortgage originations data for the 50 US states were provided by the US Department of Housing and Urban Development. These data are in current dollar terms, and are available for 1983-94. We measure real mortgage originations by dividing the current dollar data by state-level CPIs.

Population data for the US states comes from the US Department of Commerce, Bureau of the Census, and were taken from the DRI data base.

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Property-price cycles and monetary policy

Christopher Kent and Philip Lowe*

Introduction

Recent increases in equity and bond prices in many countries have renewed interest in the implications of asset-price changes for monetary policy. The reasons for this interest are clear. Rising asset prices can provide policy-makers with useful information about the likely future path of the economy; they can lead to increases in aggregate demand and inflation; and perhaps most importantly, they can sow the seeds for future financial-system problems.

This paper focuses on the last of these influences, paying particular attention to the interaction of cycles in property prices and corporate borrowing. Increases in real-estate prices generate collateral for additional loans; this stimulates credit growth and prolongs the upswing of the business cycle. However, if the price increases turn out to be unsustainable, much of this collateral can disappear, causing large losses for financial institutions and other firms. The outcome can be a pronounced and protracted slowdown in economic activity.

The fact that changes in asset prices can adversely affect the process of financial intermediation means that asset-price fluctuations can have important implications for bank supervisors and for central banks in their role as custodian of the stability of financial system. While rising asset prices might also increase expected inflation in the short run, in many cases the more important issue for central banks will be preserving the stability of the financial system. While this task is generally thought to be the responsibility of bank regulatory policy, we argue that in some circumstances monetary policy can also play a role. In particular, if the central bank can affect the probability of an asset-price bubble bursting by changing interest rates, it may be optimal to use monetary policy to influence the path of the bubble, even if it means that expected inflation deviates from the central bank's target in the short term.

Clearly, regulatory policy needs to be responsive to the destabilising effects of asset-price bubbles on the financial system. But a monetary-policy framework which is concerned with both the expected inflation rate and the variability of inflation may see monetary policy respond to asset-price bubbles in a way that also helps preserve the stability of the system. By seeking to change the path of the bubble, monetary policy may be able to simultaneously contribute to maintaining financial-system stability and to reducing the variance of inflation.

In line with the topic of this conference, this paper pays particular attention to the monetary-policy implications of asset-price changes.

The structure of the paper is as follows. Section 1 uses a simple model to analyse at a conceptual level the links between monetary policy and asset prices. It pays particular attention to the issues of how monetary policy should respond to bubbles in asset prices and how the advent of low inflation can affect the behaviour of asset prices and the response of monetary policy to changes in these prices. The important contribution of the model is that it incorporates the effect of falling

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nominal asset prices on inflation through the adverse effects that these falls have on financial intermediation.

Section 2 discusses cycles in asset prices in Australia and highlights the important link between property prices, credit and activity. We show that cycles in property prices are closely linked to cycles in credit and output while the link between equity prices and credit is much weaker. Furthermore, we suggest that property-price falls contribute to the breakdown in financial intermediation, thereby prolonging the process of recovery during downturns in activity.

We focus attention on the commercial property-price cycles of the early 1970s and late 1980s. In both cases there were substantial increases in the real price of property which were later unwound. In both cases the decline in the real price was over 50%, although in the latter period much more of the adjustment occurred through falls in nominal property prices. The effect of these cycles in property prices can be clearly seen in the ratio of credit to GDP. Despite differences in the extent of bank regulation, this ratio rose strongly during both property-price booms and then fell considerably during the downturn in property prices. We argue that the decline in credit contributed to the slow recoveries from the mid-1970s and early-1990s recessions. In contrast, the recovery from the early-1980s recession was comparatively quick, partly because the headwinds caused by declining property prices were much weaker.

Section 3 presents some econometric evidence suggesting that the property-price cycle helps to explain the business cycle and that falls in nominal property prices are particularly important. Finally, Section 4 brings together the conceptual and empirical sections of the papers by drawing some broad lessons for monetary policy.

1. Asset prices and monetary policy

In an inflation-targeting regime how should monetary policy respond to movements in asset prices? Recently, this question has received more than the usual degree of attention. Yet, the only broadly based consensus that exists is that the question is a difficult one. At one end of the debate are those who argue that asset prices should be included in the overall price index targeted by the central bank, while at the other end are those who argue that asset prices are relevant to monetary policy only in so far as they affect forecasts of future goods and services price inflation. There are also those who see a role for monetary policy in responding to asset-price movements which might threaten the stability of the financial system.

Examining all the complex interactions between asset-price movements, the macroeconomy and the health of the financial system is a difficult task and is beyond the scope of this paper. Instead we set ourselves the modest task of highlighting some relevant issues regarding the interaction of asset prices and monetary policy.

The existing theoretical literature makes two reasonably robust points. First, in assessing possible implications of asset-price changes for monetary policy, it is important to understand the source of the change in prices. Second, the prices of assets that are used as collateral for loans from financial institutions are likely to be more relevant to the setting of monetary policy than are the prices of other assets. We briefly discuss these two points and then turn to the following two questions:

- If asset price changes are not based on fundamentals what should monetary policy do?
- Does low inflation of goods and services prices affect how monetary policy should respond to changes in asset prices?

1.1 The source of the change is important

The first general point is that not all asset-price changes are the same. Prices can move for a variety of reasons and understanding those reasons is important in determining the appropriate monetary-policy response (see Smets (1997)).

In Australia's case, the most obvious example of this point is the exchange rate. As the terms of trade rise, the exchange rate tends to appreciate (see Gruen and Dwyer (1995), Smets (1997) and Tarditi (1996)). This appreciation helps reduce any inflationary impact that would otherwise be associated with the increase in the terms of trade; thus the case for an easing of monetary policy in response to an appreciation of the currency is less than clear. In contrast, if the exchange rate appreciation is not based on underlying fundamentals, there is a much stronger case for a change in monetary policy, especially if the appreciation is likely to reduce expected inflation.

Another example is the stock market. If a rise in equity prices reflects an improved outlook for corporate profits as a result of faster underlying productivity growth, the central bank's forecast of future inflation may actually fall. In contrast, if the rise in equity prices has no fundamental justification, expected inflation might be higher, particularly if aggregate demand responds to perceptions of higher wealth.

In practice, the difficult problem is accurately assessing the reason for the change in asset prices. The Sydney housing market in the late 1980s is a good example. The Sydney region had benefited from the deregulation of the financial system earlier in the decade and the increasing international integration of the Australian economy. There was a general perception that these improved "fundamentals" justified a rise in the real price of property. This perception appears to have been correct; over the four years to 1997 the average *real* price of a house in Sydney has been almost 40% higher than the average real price over the years 1983 to 1987. However, while real housing prices are now clearly higher than in the mid-1980s, they have fallen considerably from their peak in 1989. (For further details see Section 2.) While most observers thought that a real increase in property prices was justified, there were no objective criteria for determining whether the increase should be 20, 40 or 80%.

A further complication which this example illustrates is that an improvement in fundamentals often generates dynamics which cause asset prices to move by more than the fundamentals would suggest. Not only do policy-makers need to identify the source of the improved fundamentals but they also need to assess how much of a given change in prices is justified by the fundamentals. These are difficult tasks.

1.2 Property prices are important

The collateral for most loans is real estate; this means that changes in the price of real estate not only have potential effects on aggregate demand, but also on the health of the financial system. If nominal property prices actually fall, and if financial institutions have made loans with relatively high loan-to-valuation ratios, the underlying collateral may be insufficient to match the face value of the loan. The result can be substantial losses by the financial system, which can have adverse effects on the future availability and cost of intermediated finance.

Changes in property prices can affect the balance sheets of corporations, as well as financial institutions. A fall in prices reduces the net value of firms and, due to imperfections in credit markets, makes it more difficult to attract intermediated finance for a given investment project (see Bernanke and Gertler (1990), Gertler (1992), Kiyotaki and Moore (1995) and Lowe and Rohling (1993)). As a result, a financial accelerator acts to amplify any downturn in economic activity (Fisher (1933)). To some extent this effect is likely to work regardless of whether the fall in property prices is in real terms or nominal terms; falling real property prices are likely to constrain access to intermediated finance, even if nominal prices are rising. In contrast, a given fall in real property prices is much more likely to generate the type of adverse financial-system consequences discussed above if the fall occurs in a low-inflation environment, with *nominal* asset prices actually falling. We return to this issue in Section 1.4.

In Australia, relatively little lending by financial intermediaries has been secured against equities. Hence booms and busts in equity prices need not have the same direct implications for the balance sheets of financial institutions that changes in property prices have had. Nevertheless, movements in equity prices can still affect the stability of the financial system. As the experience of the United States in 1987 shows, a major fall in equity prices can create problems in the payments system, with potentially quite large adverse consequences. Further, if a share market crash leads to a severe contraction in aggregate demand, borrowers may find themselves unable to repay their loans.[1] As share ownership becomes more widespread, the aggregate demand effects of changes in equity prices may become more pronounced.[2] Continued financial innovation may also see the growth of lending secured against equities which would add to the exposure of balance sheets of financial institutions to changes in equity prices. Such a change in the pattern of financial intermediation would increase the relevance of stock prices for monetary policy.

2.3 Monetary policy and bubbles: a simple model

In this subsection we take it as given that the increase in asset prices is not justified by underlying fundamentals and that the central bank knows this. We loosely refer to increases of this kind as bubbles. To discuss their implications for monetary policy we use the following simple model of the economy:

$$\pi_t = \alpha A_t + \beta D_t \Delta A_t - R_{t-1} \tag{1}$$

where π_t is the deviation of inflation from the central bank's target, A_t is the deviation of the asset price from its fundamental value, R_t is the deviation of the policy interest rate from its neutral level and D_t is a dummy variable which takes a value of 1 if the asset price has fallen, and 0 otherwise. While the target variable is taken to be inflation, the following analysis would apply equally if the target were the deviation of output from its potential level. Note that monetary policy is assumed to affect the inflation rate with a lag of one period.

In this model, the effect of a change in the asset price on inflation is asymmetric. This asymmetry arises from the effect that declines in nominal prices have on the health of the financial system.

If the asset price is originally at its fundamental value and then *increases*, economic activity expands, and inflation increases. A higher asset price creates perceptions of greater wealth, leading to an expansion of aggregate demand and rising inflationary pressures, as there is no improvement on the supply side of the economy; the larger α is the larger is the effect. However, when the bubble bursts and the asset price returns to its fundamental value, not only are the stimulatory effects of the higher asset price withdrawn, but there are additional contractionary effects. These effects arise from the balance-sheet problems discussed above and act as a disinflationary force on the economy; the larger β is, the more powerful is this force. The experience of the past decade suggests that β is likely to be larger for property prices than for equity prices.

When the asset price is above its fundamental value, there is some probability p that it will return to its fundamental value next period. This probability is assumed to be a function of the

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1 Equity prices can influence aggregate demand through wealth effects and through their effect on the cost of capital (see Anderson and Subbaraman (1996) and de Roos and Russell (1996)).

2 The 1997 Australian Sharemarket Survey finds that 34% of the Australian adult population has direct or indirect share holdings. This percentage would be considerably higher if account was taken of employer-funded pension schemes.

deviation of the interest rate from its neutral level. Higher interest rates increase the debt servicing on purchases of the asset and make a downturn in the business cycle more likely; these changes increase the probability that the bubble will collapse. In addition, a decision by the central bank to increase interest rates could be accompanied by remarks regarding the high level of asset prices and the unsustainability of current trends; such remarks might also make the continuation of the bubble less likely.

We model the relationship between today's interest rate and the probability of the bubble collapsing in the next period as follows:

$$p_{t+1} = \phi + \varphi R_t \tag{2}$$

The larger φ is, the larger is the effect of the interest rate on the probability of the bubble collapsing. If the bubble does not burst it is assumed to grow at rate g^* so that:[3]

$$A_{t+1} = gA_t \tag{3}$$

where $g = 1 + g^*$.

Further, we assume that the economy operates for three periods; that a bubble emerges in the first period; and that if the bubble collapses in the second period it does not subsequently re-emerge in the third period.[4] While this last assumption is crucial in deriving the results below, it is not important that it holds exactly; what is important is that that the probability of the bubble re-emerging is relatively small, at least in the near term. An examination of real-world bubbles suggests that this requirement is not a stringent one (see Section 2).

Having observed that a bubble has emerged in period 1, the task for the central bank is to minimise the (undiscounted) sum of the squared deviations of inflation from its target. Since the central bank cannot affect the current rate of inflation, this amounts to minimising:

$$E_1(\pi_2^2) + E_1(\pi_3^2) \tag{4}$$

where E denotes the expected value and subscripts refer to time periods. This objective function assumes that the central bank is not only concerned about the expected value of inflation but also the variability of inflation; this latter concern may reflect risk aversion or the real costs associated with variability in the inflation rate.

We solve this problem recursively, solving first for the two possible interest rates in period 2; one for the case in which the bubble bursts in period 2 and one for the case in which the bubble does not burst. Using these solutions we then solve for the optimal interest rate in period 1.

The solutions are analytically quite complicated since non-linearities are introduced by making the probabilities of various outcomes a function of the interest rate.[5] Thus, rather than present

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3 We have assumed that the growth rate of a surviving bubble (g*) is independent of the probability of the bubble collapsing. However, rational expectations require that the growth rate of the bubble increases as the probability of the bubble collapsing increases. In addition, under rational expectations, an increase in interest rates (which increases the probability of collapse of the bubble), also would lead to a discrete fall in the current size of the bubble. While our model is not compatible with rational expectations, we regard it as realistic to assume that the central bank takes as exogenous the growth rate of a surviving bubble.

4 Three periods is the minimum number needed to generate interesting results.

5 The analytical solutions were derived with Mathematica Version 3.0.

the algebraic solution we discuss the solutions for a few different parameter sets. A summary of the results is presented in Table 1.

Table 1

Model results

Parameter/ Variable	Definition	Case 1			Case 2	
		1a	1b	1c	2a	2b
g	Growth rate of bubble	2.0	2.0	3.0	2.0	2.0
ϕ	Exogenous probability of collapse	0.5	0.5	0.5	0.2	0.2
φ	Effect of interest rates on probability of collapse	0.0	0.2	0.2	0.0	0.2
β	Cost of bubble collapse	2.0	2.0	2.0	2.0	2.0
R_1	Interest rate in period 1	0.0	0.6	2.1	1.2	1.0
p_1	Probability of collapse in period 1	0.5	0.62	0.92	0.2	0.4
$E_1(\pi_2)$	Expected inflation in period 2	0.0	-1.1	-3.7	0.0	-0.6

Note: The inflationary effect of the bubble, α, and the initial size of the bubble, A_1, are always set equal to 1.

In the following examples we set $\alpha = 1$, $\beta = 2$ and the size of the bubble in the first period (A_1) equal to 1. These parameter values imply that a collapse in asset prices has a large negative effect on output relative to the expansionary effect of the bubble. Initially we set $g = 2$ (if the bubble survives, it doubles in size each period) and $\phi = 0.5$ (if the interest rate is at its neutral level, the probability of the bubble bursting next period is 0.5).

Now suppose that changing the interest rate has *no* effect on the probability of the bubble bursting ($\varphi = 0$); this is case 1a in Table 1. In this example the optimal policy is to leave the interest rate unchanged (at the neutral rate); if the bubble continues next period, inflation will equal +2, if it collapses inflation will equal -2. Given that these two events have the same probability of occurring, the expected loss from the bubble crashing is equal to the expected loss from the bubble continuing. The expected variance of inflation in period 2 is 4 and this cannot be reduced by changing the interest rate.

In this example, monetary policy cannot affect the probability of the bubble bursting and so possible events beyond the standard transmission lag (one period) have no bearing on the current setting of policy. Of course, possible events next period are important; if we increase the growth rate of the bubble, reduce the probability of the bubble collapsing, or reduce the costs associated with a bubble collapse, interest rates should be increased, rather than held constant.

Now instead, suppose that the central bank can affect the probability of the bubble bursting (say $\varphi = 0.2$). In this case (1b in Table 1) the optimal policy is to raise the interest rate by 0.6 above the neutral rate.[6] The higher interest rate increases the probability of the bubble collapsing from

[6] The units are not important here, as they could be rescaled by introducing a parameter on the interest rate in Equation (1).

0.5 to 0.62. The effect of this policy is to cause expected inflation to be below target in the second period.

Where does this result come from? Monetary policy has two effects in this model: one standard and one non-standard. First, if the central bank takes the probability of the bubble collapsing as fixed, it faces the usual problem of making decisions under uncertainty. The bank solves this problem by determining the possible levels of inflation in the second period and then, taking the relevant probabilities as given, solves for the interest rate that minimises the expected variance of inflation around the target. The expected rate of inflation is equal to the bank's target.

The second, and non-standard, effect of monetary policy is on the probability of the bubble collapsing. If the central bank can affect this probability, it needs to take into account not only possible outcomes in period 2, but also possible outcomes in period 3. While the standard transmission lag is only one period, monetary policy can affect the probability of events occurring in subsequent periods. If policy can burst the bubble in period 2, the probability of the bubble bursting in period 3 is reduced to zero. In this example, it makes sense for the central bank to raise the interest rate to increase the probability of the bubble bursting in period 2 and thereby reduce the probability of even more extreme outcomes in period 3.

As the growth rate of the bubble increases, the optimal interest rate increases at an increasing rate. A higher growth rate of the bubble increases the variance of possible outcomes in each of the following periods, with the effect being substantially larger for the third period. As a result, the pay-off to bursting the bubble in period 2 rises. For example, if g equals 3 rather than 2 (so that the bubble triples in size each period, rather than doubles), the optimal interest rate in the first period is 2.1 above the neutral level and the probability of the bubble collapsing increases to 0.92 (case 1c in Table 1). In a sense, the large increase in the interest rate amounts to a strategic attack on the bubble. Policy is tightened to such an extent that it is almost certain the bubble will break. The policy-maker knows that this will be quite contractionary; not only will growth be retarded by the lagged effect of the high interest rate, but it will also be adversely affected by the flow-on effects of the fall in asset prices. Yet failing to increase the interest rate results in a much higher probability of a larger crash at some later point.

In case 1a above, optimal policy involves maintaining the interest rate at zero when the probability of the bubble bursting is exogenous, but increasing the interest rate when the probability is endogenous. We now consider a case in which the optimal policy is to increase the interest rate when the probability of the bubble bursting is exogenous. To do this we reduce ϕ (the exogenous probability of collapse) from 0.5 to 0.2, while keeping all other parameters unchanged (case 2a). With a reduced probability of the bubble bursting next period, it is more likely that the inflation will be above target in the second period and so the interest rate in period 1 is higher than it was in the earlier case (1.2 compared to 0). If we now make the probability of collapse endogenous ($\varphi = 0.2$), the optimal policy again sees interest rates increasing (to 1.0; case 2b), but the increase is smaller than the increase when the probability of collapse is not influenced by monetary policy. Once again, the higher interest rates increase the probability of the bubble bursting in period 2 from 0.2 to 0.4.

In this example, the higher interest rate needed to counteract the expansionary effect of the asset-price bubble also increases the probability of the bubble collapsing; this amplifies the expected effect of the tightening of policy. This allows the central bank to increase the interest rate by a smaller amount than would otherwise be the case.

This result comes partly from the way we have modelled the probability of the bubble collapsing (Equation 2). If this probability is a function not of the deviation of the interest rate from the neutral level, but of the difference between the interest rate and the rate needed to offset the standard expected demand effects of the bubble (1.2 in the above example), then monetary policy would always be tightened by more than the standard analysis would suggest.

The above examples highlight a few general points:

- If the probability of a bubble bursting is related to the current level of interest rates, the central bank may wish to raise interest rates to increase the probability of the bubble breaking, even if an increase in interest rates cannot be justified on more conventional grounds. Such a policy would cause an expected slowdown or contraction in activity with inflation falling below target; the gain would be a reduction in the variance of possible outcomes in the following period.
- The monetary authorities need to be concerned with possible outcomes beyond the period of the normal transmission lag. If the probability of a bubble collapsing is endogenous, the expected path of the economy beyond the control lag is affected by today's interest-rate decision.
- If interest rates need to be increased to offset the expansionary effect of a bubble, the size of the optimal interest-rate increase may be smaller if the probability is endogenous rather than exogenous. By adding what amounts to an additional transmission channel, the size of the optimal response may decline.

Finally, in the above model we have assumed that the parameters α and β are fixed and independent of monetary policy. If these parameters could be lowered, the need for monetary policy to respond to changes in asset prices would be reduced. In the limit, if they were both zero, no monetary response to asset-price bubbles would be required. Alternatively, if β could be reduced to zero (so that falls in asset prices did not have implications for the health of the financial system) monetary policy could just be concerned with the standard demand effects of asset-price changes. However, while it might be desirable to reduce these parameters, there is probably little that monetary policy can do in this regard.

One alternative might be to use financial regulation as a second instrument. For example, if the supervisory authorities thought that a serious bubble had emerged, they might be able to reduce α and β (or g) by raising capital requirements on property loans, and/or by imposing low maximum loan-to-valuation ratios. By limiting the extent to which financial institutions could make loans collateralised against property, such policies might reduce the contractionary effects when the bubble finally bursts; they might also increase the probability of the bubble bursting. In general, if such a policy option is available it may be preferable to using interest rates; a primary advantage would be that when the bubble bursts, the economy would not also be dealing with the lagged effects of high interest rates.

Notwithstanding this advantage, the use of financial regulation to affect asset prices is not without considerable difficulties. Foremost amongst these is that imposing regulations on the banking sector may simply induce a shift towards institutions that are more lightly regulated. In a sense this was the experience in Australia in the early 1970s. Large increases in property prices were accompanied by rapid growth of non-bank financial institutions as well as other vehicles for investing in property. When the property crash came, losses were concentrated in these institutions rather than in the banks, but there were still considerable contractionary effects on the economy (see the discussion in Section 2). Another difficulty is that imposing and lifting regulation may affect the efficiency of the financial system.

1.4 Implications of low inflation

The 1990s have seen a remarkable convergence amongst OECD countries in goods and services price inflation, with most countries now having inflation rates of 3% or less. For those countries with previously high rates of inflation, the advent of stability of goods and services prices has a number of implications for asset prices and the interaction of asset prices and monetary policy. We briefly discuss four of these.

First, low inflation should reduce the likelihood of a bubble originating. In an environment in which inflation is high and variable, property acts as a hedge against inflation and there are also often substantial tax advantages to investing in property. In contrast, in an environment with low and stable inflation, the incentive to purchase property is diminished. These changed incentives should reduce the likelihood of speculative increases in property prices.

Second, and perhaps most importantly, if an asset-price bubble does occur, a low-inflation environment makes it more likely that the inevitable correction in *real* asset prices will occur through a decline in *nominal* asset prices. The Australian experience is again instructive. After the property-price boom of the early 1970s, real property prices declined significantly; while this involved some fall in nominal prices, the high background rate of goods and services price inflation accounted for much of the adjustment in real prices. In contrast, in the low-inflation 1990s, relatively more of the adjustment in real property prices took place through the nominal price of property falling (see Section 2). Given that in a low-inflation environment, financial institutions are less likely to realise the full nominal value of collateral if borrowers fail to repay their loans, loan-to-valuation ratios should decline. If this does not occur, a correction in real property prices, through declining nominal prices, can significantly exacerbate the business cycle. If this is the case, it makes it more important that monetary policy responds relatively early in the life of the bubble.

Third, lower nominal interest rates associated with lower inflation reduce the liquidity constraints that previously restricted the access of some borrowers to intermediated finance. In Australia, it is not uncommon for financial institutions to determine the maximum amount an individual can borrow for the purchase of a house by calculating the size of the loan which would generate initial repayments equal to 30% of the individual's income. As a result, lower interest rates mean larger loans and more individuals qualifying for a housing loan (see Stevens (1997)). This is likely to increase the dispersion of house prices in any local market, and may affect the average price as well. Upward pressure on prices might also occur if reduced liquidity constraints lead to an increased rate of household formation.

Finally, average real interest rates are likely to be lower in a period of sustained low inflation than they are in the period of disinflation. This reflects not only the easing of the previously restrictive monetary policy, but also lower risk premia associated with low and stable rates of inflation. Lower expected real interest rates should lead to a rise in real asset prices. In addition, a number of authors have argued that low inflation reduces the "equity risk premium", and that, as a result, real equity prices should increase as low inflation becomes entrenched.[7] These effects can be quite substantial. If one uses the discounted dividend equity-valuation model, a one percentage point decline in the real interest rate, or the equity premium, generates an increase in the equity price of 25% if the initial dividend yield is 4%.

In summary, a move from moderate to low inflation rates might justify some fundamental increase in real asset prices, although as usual, determining the extent of this effect is difficult, and improving fundamentals can themselves be conducive to generating bubbles. While, in the medium term, low and stable inflation should probably reduce the likelihood of bubbles occurring, if a bubble does occur, there are perhaps stronger implications for the health of the financial system. As a result, the returns to early action to increase the probability of the bubble collapsing may be higher.

2. Cycles and bubbles in Australian asset prices

In this section of the paper we study cycles in the prices of Australian equities, commercial property and residential dwellings over the past three decades. We identify three broad cycles. The first begins from the late 1960s for equity prices and in the early 1970s for property prices, the second cycle starts in the late 1970s and the third cycle starts after 1985. Of the three cycles, the early-1970s and late-1980s cycles are clearly the more significant.

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7 See Modigliani and Cohn (1979) for a theoretical justification of this argument and Blanchard (1993) and Kortian (1997) for empirical evidence.

Figure 1
Australian nominal asset price cycles

September 1979 = 100

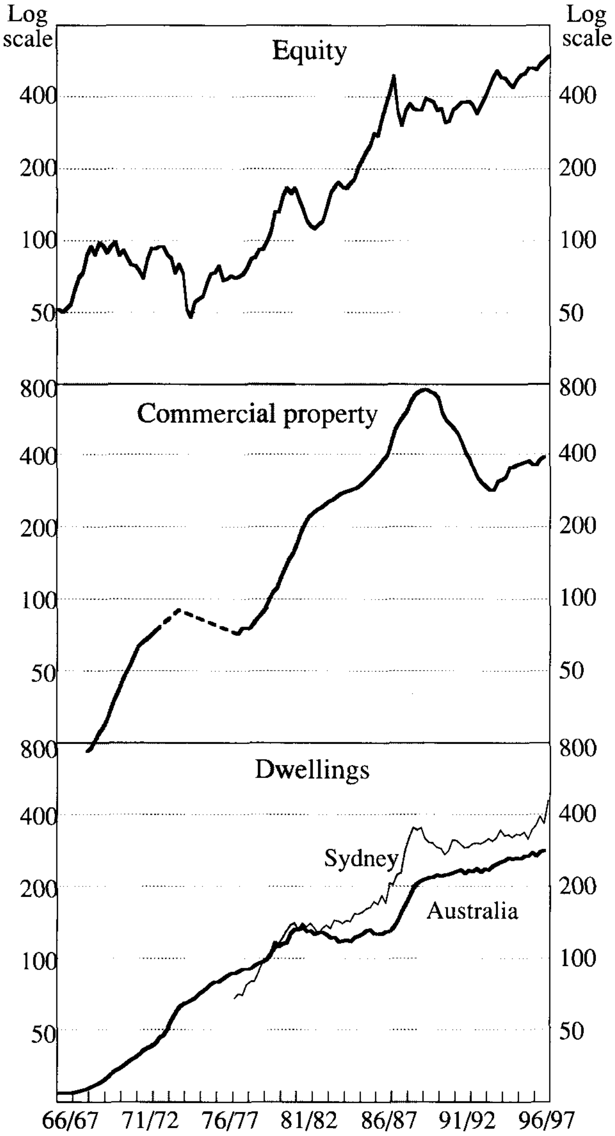
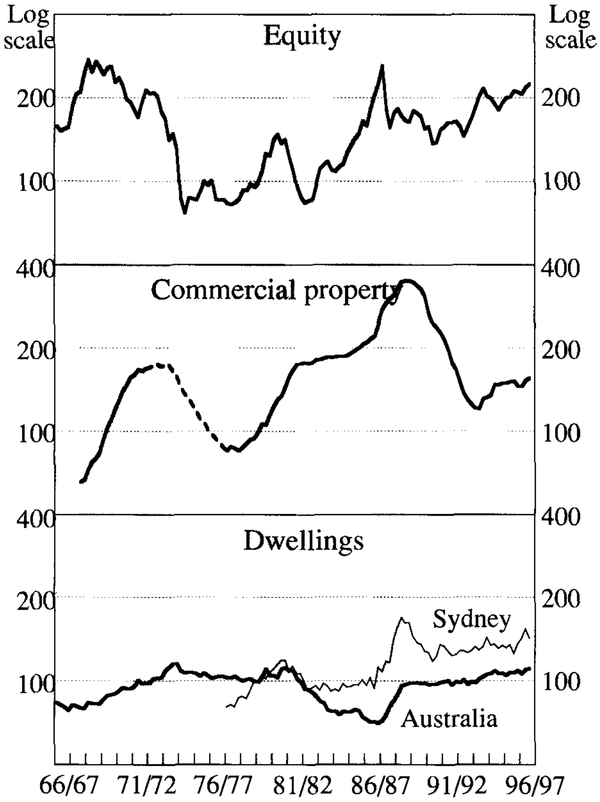


Figure 2
Australian real asset price cycles

September 1979 = 100



Note: Shading indicates the upturn phase of each cycle. Broken line segment is an estimate (see the Appendix).

Figure 1 shows these three cycles in nominal terms and Figure 2 shows them in real terms (the upswings of the cycles are shaded). Table 2 provides details of the precise size, duration and growth rates of each phase of each of these cycles - upturns, downturns and periods in between. Details of the data and the dates of each phase of the cycle are recorded in the Appendix. Our goal is to indicate the broad patterns in asset prices rather than to pick the exact dates of each phase in the cycle.[8]

Between March 1968 and March 1997, nominal equity prices increased eight-fold, commercial-property prices seventeen-fold, and dwelling prices ten-fold. The percentage increases in

[8] It could be argued that the start of the most recent cycle in commercial property prices was a couple of years earlier than we have shown. However, the same could be said of equity prices. These sorts of minor adjustments would not change our main observations nor the substance of our arguments.

real terms over the same period were 7, 137 and 39% for equity, commercial property and dwelling prices respectively. The largest swings through cycles were in commercial property, whereas movements in dwelling prices were comparatively muted (Table 2).

Table 2

Cycles in real asset prices

Average annualised quarterly growth rate (in per cent), duration (in quarters)and size (percentage change)

	Equity			**Commercial property**			**Dwellings**		
	Growth rate	**Duration**	**Size**	**Growth rate**	**Duration**	**Size**	**Growth rate**	**Duration**	**Size**
Early-1970s cycle									
(in between)	-	-	-	-	-	-	3	20	15
Upturn	20	13	71	19	23	164	7	11	21
Downturn	-21	19	-71	-17	15	-51	-3	21	-14
(in between)	4	13	9	0	3	0	-	-	-
Late-1970s/ Early-1980s cycle									
Upturn	18	13	69	21	15	106	4	9	10
Downturn	-34	5	-41	2	12	7	-11	12	-31
(in between)	15	8	30	9	8	19	-3	10	-7
Late-1980s cycle									
Upturn	35	12	141	20	10	57	20	7	38
Downturn	-9	21	-45	-23	16	-65	0	16	0
(in between)	11	17	55	8	14	29	4	16	15

Note: The "in between" portion of the early-1970s cycle for dwellings occurs prior to the upturn. The 1970s downturn for dwellings ends with the late-1970s upturn.

The most recent cycle, which followed the deregulation of the financial system, was in many ways more pronounced than earlier cycles. For equity and dwelling prices, the late-1980s upturn was the most rapid and largest of the three cycles. Also, the 1990s downturn in commercial-property prices was the largest and most rapid (in both nominal and real terms). Despite the significant rise in commercial-property prices in the late 1980s, the percentage increase in *real* prices was actually smaller than in the early-1970s cycle, and the duration of the cycle was shorter. These observations remain true even after we control for differences in the level of real activity across the three cycles - that is, by examining the ratio of nominal asset prices to nominal GDP.

A pattern emerges when we examine the timing of cycles across different asset classes. Namely, equity-price cycles lead property-price cycles; commercial-property and dwelling prices turn down together. The fact that equity prices move before property prices could reflect two things. First, although property prices are set in forward-looking markets, equity markets are much more liquid; therefore, we expect equity prices to move first and fastest. The relatively volatile behaviour of equity prices is apparent in Figures 1 and 2. Second, as equity markets turn down, investors seek to move out of equities and into other assets, including property, which at the time would be seen as both safer and more likely to provide better returns. Daly (1982) argues that this switching between asset types was partially responsible for the early-1970s upturn in property prices, and the same may also be true of events following the October 1987 share-market crash.[9] Finally, the observation that commercial property and dwelling prices turn down together may simply reflect the fact that the downturns in both of these markets are closely associated with recessions (see below).

We do not formally attempt to identify asset-price bubbles. Instead we simply assert that *ex post*, large downward corrections in real asset prices occurring immediately after a period of rapid real increases indicate that the boom did have a substantial bubble component. Using this criterion a number of cycles stand out as having a substantial bubble component; in particular, the cycles in commercial-property prices in the early 1970s and late 1980s, as well as the cycles in both equity and Sydney dwelling prices during the mid to late 1980s. In each of these cases, large and rapid real increases were promptly followed by rapid real falls in asset price (see Figure 2 and Table 2).

Figure 2 suggests that bubbles do not restart once the real price has started to fall. There is a considerable interval between bubbles (and between price cycles more generally). This is consistent with the key assumption of the model in Section 1 that once bubbles burst they stay burst (for a reasonably long period).

The more general point to make here is that bubbles are often closely linked to the business cycle (Blundell-Wignall and Bullock (1992)). The expansion phase of a cycle justifies some fundamental increase in the real asset price. If this expansion is accompanied by other factors that support a real increase in asset prices (such as financial deregulation), then further real asset price increases will occur. However, it becomes difficult to determine to what extent these real asset price rises are justified by fundamentals and to what extent they constitute the beginning of a speculative bubble. There is no evidence of bubbles occurring during times of weak fundamentals (like recessions). Similarly, the collapse of a bubble is more likely when fundamentals are expected to worsen.

The relationships between asset prices, the business cycle and the cycle in credit are shown graphically in Figure 3.[10] Three patterns emerge. First, a recession coincides with each fall (or major slowing) in real property prices. Recessions in 1974, 1982/83 and 1990/91 occurred within only a few quarters of the end of the upturn phases of the commercial-property and dwelling price cycles.[11] This close link is not apparent for equity prices; most notably, the fall in real equity prices in 1987 was not associated with any significant slowdown in activity.

Second, credit cycles are closely linked to cycles in property prices. The ratio of credit to GDP has been steadily rising through time; this is consistent with financial deregulation and with the positive correlation between GDP per head and the extent of financial intermediation. The largest deviations from this trend increase in credit coincide with the early-1970s and late-1980s property-price cycles. Credit rose more rapidly during the substantial upturns in commercial-property prices in these two cycles, but then fell (as a share of GDP) during the downturn in the property market. The importance of property prices in influencing developments in credit is underlined by a comparison of these two cycles with the experience in the early 1980s. Despite a very severe recession in 1982/83,

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9 Kent and Scott (1991) show that there was a marked increase in office construction and investment by the finance sector at about this time. It is interesting to note that the October 1997 equity price collapse did nothing to dampen investment by the finance sector.

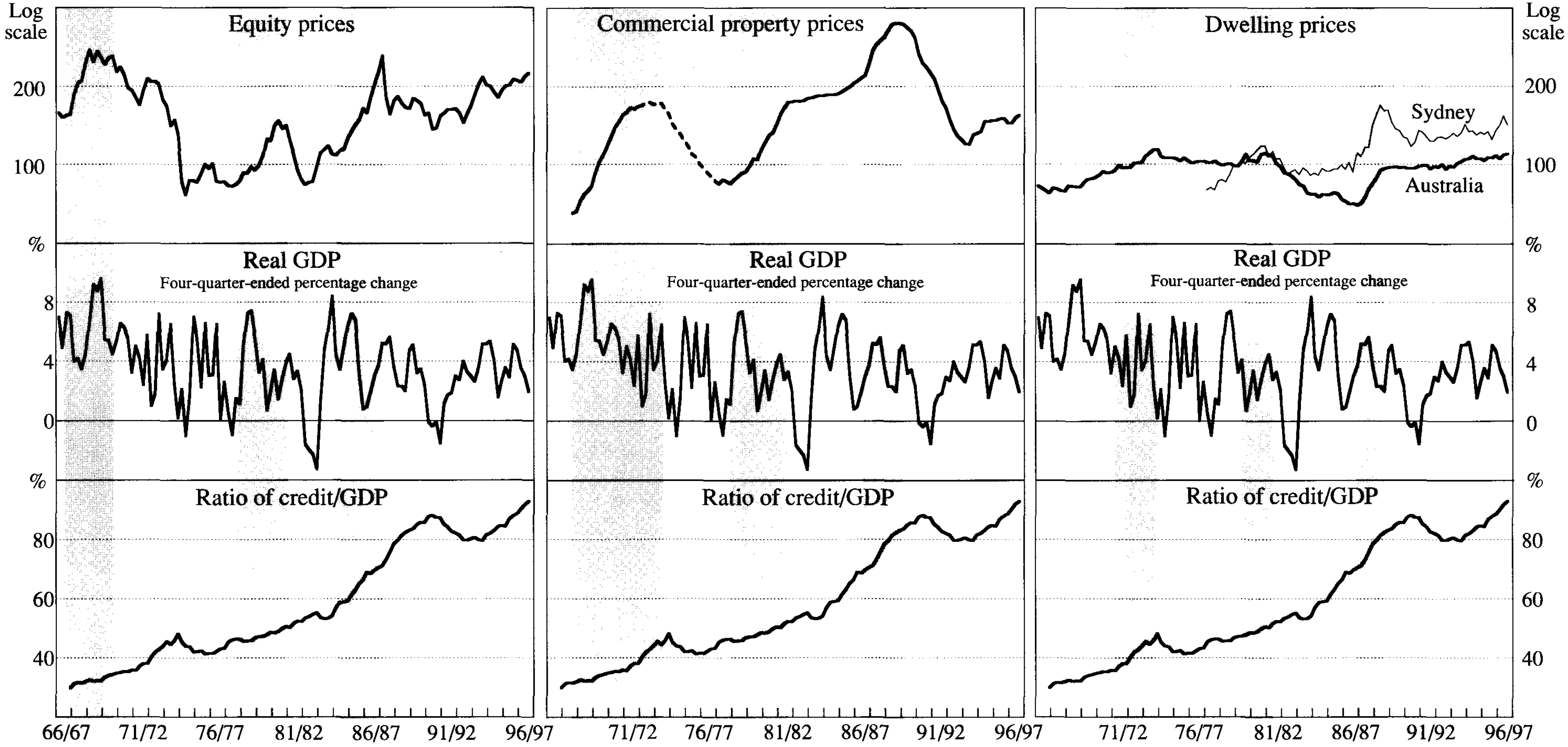
10 For a comprehensive review of these relationships across a wider range of countries see Borio, Kennedy and Prowse (1994).

11 Of course, this does not imply causation running from asset price falls to recessions. The linkages we discussed above run both ways and asset prices are forward looking.

Figure 3

Real asset prices, real GDP growth and credit

Asset prices: September 1979 = 100



Note: Shading indicates the upturn phase of each cycle. Broken line segment is an estimate (see the Appendix).

the ratio of credit to GDP fell only marginally, and quickly recovered its earlier level. In contrast, in the other two cases it took over 5 years for the ratio of credit to GDP to recover its earlier peak. A plausible explanation for this difference is that the muted property-price cycle in the late 1970s/early 1980s led to a muted cycle in the ratio of credit to GDP. In other words, falling property prices in the early 1970s and early 1990s led to greater falls in credit and more protracted recessions (see below) because falls in the value of collateral caused falls in the degree of financial intermediation.

Third, in contrast, equity-price cycles appear to be unrelated to cycles in credit. In fact, credit continued to grow strongly during each downturn in equity prices.

The early-1970s and late-1980s cycles in commercial-property prices are particularly interesting because they occurred in very different financial and inflation environments and yet they were surprisingly similar in many ways.

Daly (1982) highlights a number of factors that led to, or sustained, the boom in property prices in the early 1970s. These factors include the increasing internationalisation of the Australian economy and the growth of non-bank financial institutions (in part through an influx of foreign merchant banks). There was a general increase in the demand for office space which was fuelled by the minerals boom of the late 1960s, the subsequent development of Sydney as a major financial centre and generally favourable business expectations.[12] Relatively easy monetary conditions also played a role.

In the early 1970s there were tentative steps towards deregulation of the financial system, with quantitative controls on lending being suspended in 1971 and the deregulation of some bank lending rates in 1972 (see Grenville (1991)). Despite these changes, the banking system remained heavily regulated; there were controls on banks' deposit and lending rates and on the composition of banks' balance sheets. These controls meant that banks were constrained in their ability to make additional loans on the back of higher property prices. This created an incentive for other financial institutions to provide real estate loans and, as a result, there was extremely rapid growth of the less-heavily regulated non-bank financial institutions; these included building societies, finance companies (some of which were bank owned) and merchant banks (see Figure 4). In addition, large capital inflows were channelled directly into commercial property, either through development companies (domestic and foreign) or through finance companies.

Monetary policy was relatively loose in the early years of the 1970s; a strong external sector saw a current account surplus and a rapidly increasing money supply. Monetary policy was not tightened until mid 1973, almost at the same time that the property-price cycle reached its peak. In April of 1973, reserve requirements were increased; in July and August, quantitative controls were again placed on bank lending;[13] and in September, (regulated) bank interest rates were raised.

Property prices peaked and turned towards the end of 1973. The initial casualties were the property developers. This led to further falls in prices and the eventual collapse of some large finance companies; in a number of cases, bond holders in these companies experienced substantial losses. In other cases, foreign-bank parents supported their Australian subsidiaries. The fact that the core banking system did not engage in the same degree of property lending as the non-bank financial institutions meant that they were less affected by the decline in property prices. However, they were not insulated completely, as some bank-owned finance companies incurred substantial losses. The most notable example was the Bank of Adelaide which was eventually merged with the ANZ Bank (albeit not until 1979).

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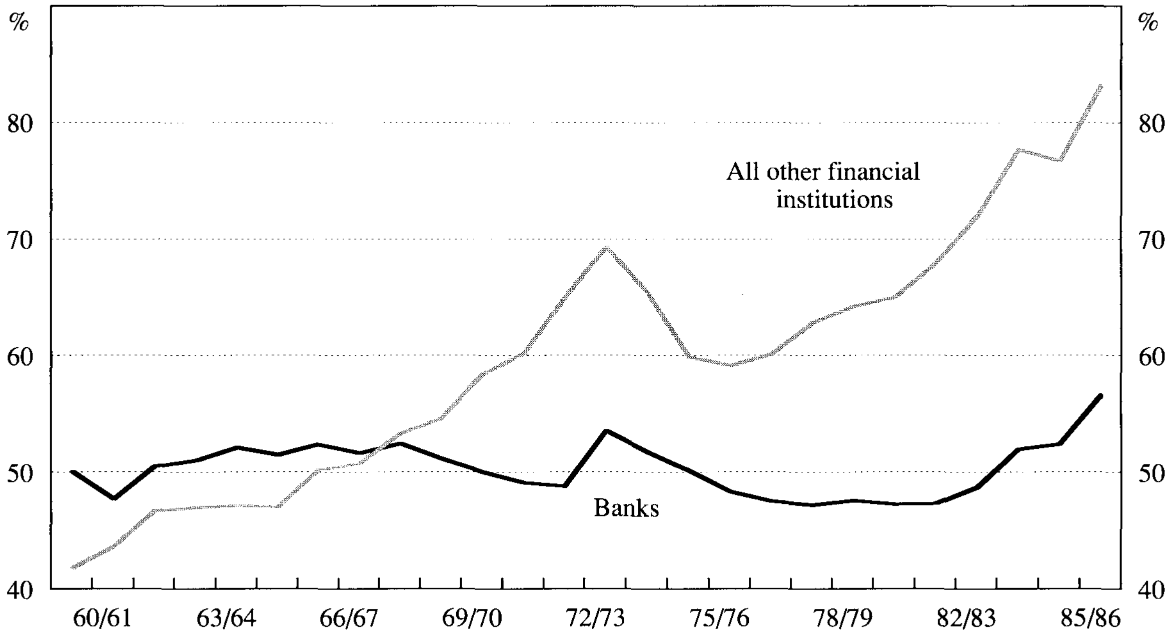
[12] By 1976 Sydney was ranked as the ninth largest international investment centre in the world, a long way ahead of Melbourne (ranked twenty-ninth) which had traditionally been Australia's financial centre (Davis (1976), p. 28).

[13] Trading banks were "requested" to achieve an appreciably lower level of new lending and savings banks were asked to limit their total housing loan approvals in the half year to December to not more than twice the level in the June 1973 quarter (Reserve Bank of Australia (1974)).

Figure 4

Total assets of financial institutions

As a percentage of GDP



The commercial property price cycle in the late 1980s shares many of the characteristics of the early-1970s cycle. The precursors were familiar: namely, financial liberalisation and increasing internationalisation (Blundell-Wignall and Bullock (1992)). In the late 1980s, however, another factor was also at work; the previous decade or so of high inflation had made it clear to people that there were considerable tax advantages to purchasing property in a high-inflation environment. This added to the demand for real estate. Another familiar characteristic in the late-1980s episode was the substantial increase in the ratio of credit to GDP.[14] While the *percentage point* increase in this ratio was certainly larger in the late 1980s, the *percentage* change in this ratio was actually higher in the 1970s.[15] A third common characteristic was the large losses by financial institutions following the boom.

A key difference between the two cycles is that in the late-1980s cycle, banks were not as restricted in their lending activities. While some lending again took place through the banks' non-bank subsidiaries, substantially more of the ultimate impact was felt in bank profitability (see Figure 5). It is difficult to pinpoint whether the decline in credit was in response to this fall in profitability, or simply reflected reduced demand for credit by firms. In all probability both explanations play a role. Falling asset prices exacerbated the already high debt-to-equity ratios of many companies, with the result that investment plans were delayed while corporate balance sheets were repaired. The demand for credit was probably further weakened by the increase in the banks' lending rates relative to official interest rates (second panel of Figure 5). Banks also adopted more cautious lending practices as they came to grips with falling property prices.

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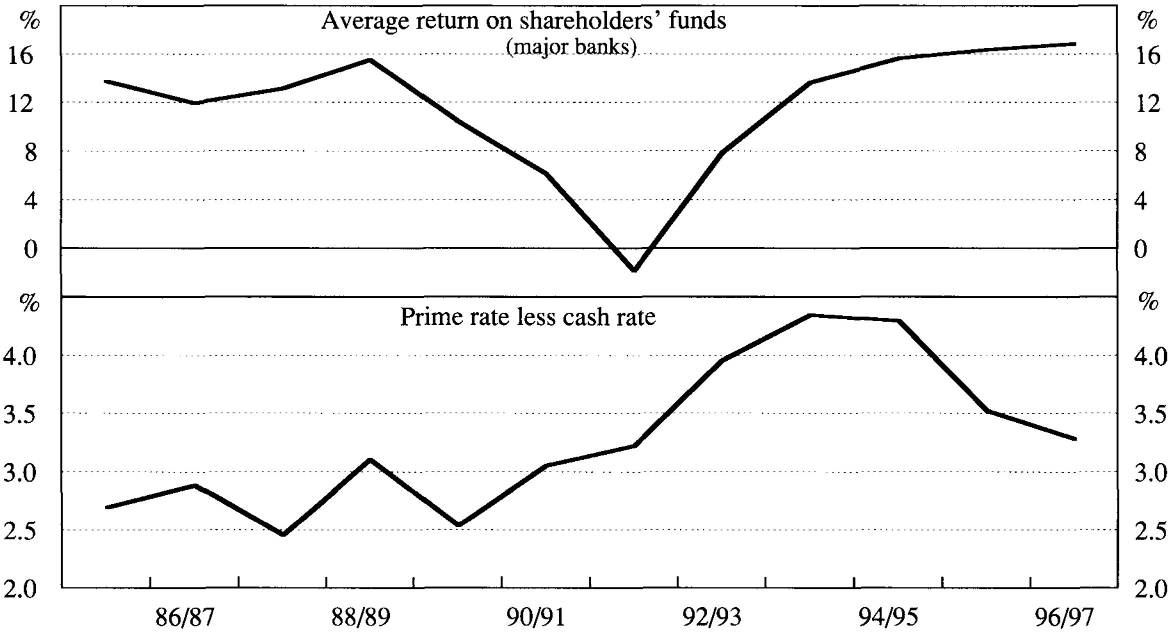
[14] For further details on the credit cycle in the late 1980s see Macfarlane (1989) and (1990).

[15] In the five years to June 1974 the credit to GDP ratio increased by 16 percentage points to 48% (which was an increase of 50%). In the five years to December 1990 the ratio increased by 25 percentage points to 89% (which was an increase of 40%).

Figure 5

Returns on bank shareholder funds and bank margins



Another difference is the way that monetary policy operated. In the early episode, the property-price cycle turned at about the same time that monetary policy was tightened. In the late-1980s episode, a relatively long period of tight monetary policy preceded the turnaround in property prices; interest rates were first increased in April 1988 but the property-price cycle did not turn until late in 1989. During this period, domestic demand was growing quite strongly and inflationary pressures were building. While the high interest rates were designed to deal with these pressures, the rapidly rising asset prices were also of concern. Macfarlane (1989) argued that high inflation had led people to conclude that purchasing assets was the road to increased wealth, and that this was leading to a severe misallocation of resources. While asset prices were not a target of monetary policy, the combination of rising real asset prices and high rates of inflation was thought to be untenable (see Macfarlane (1991) and Grenville (1997) for a more detailed discussion).

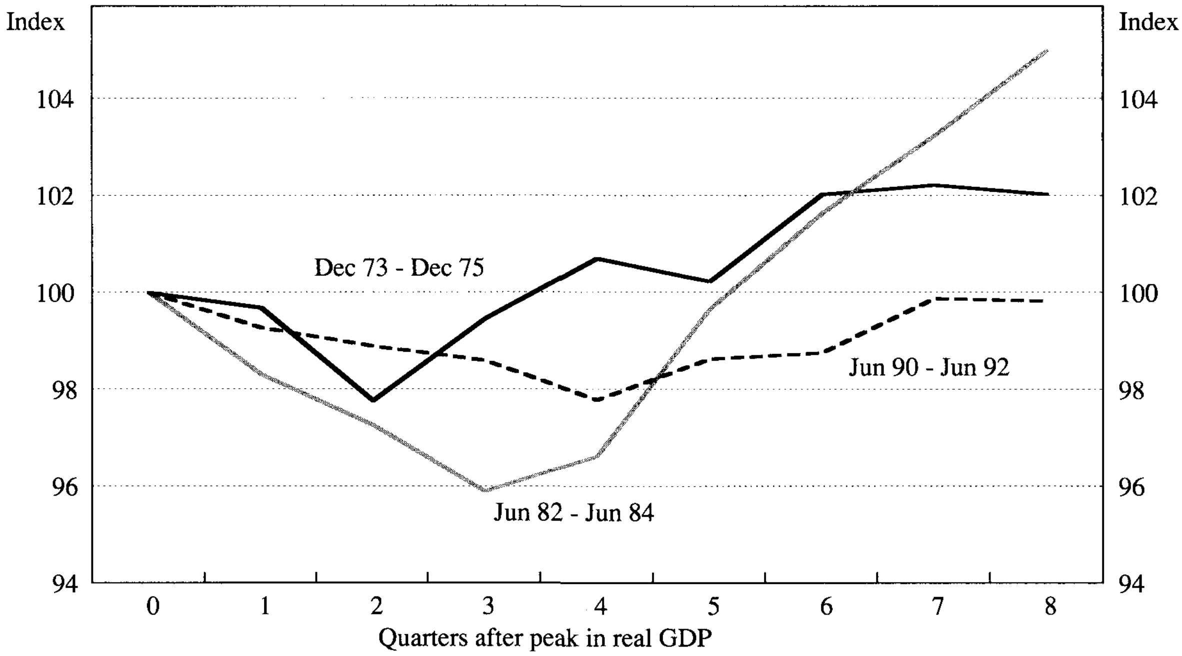
In neither episode did monetary policy set out explicitly to break the bubble in property prices, although, in both cases, some effect on property prices was expected. In the 1970s episode, quantity controls on bank lending, and slower growth in the money supply resulting from a deterioration in the balance of payments, appear to have had a relatively rapid effect on property prices, although the boom had probably been running out of steam at the time that monetary conditions were tightened. In the 1980s episode, the entrenched inflation culture made affecting property prices more difficult. With large increases in property prices expected, real interest rates (calculated using expected property-price inflation) were very low, despite more-conventionally measured real interest rates being quite high. Only sustained high financing costs and an increased probability of a significant slowdown in the pace of economic activity brought the boom to an end. As a consequence, when the decline in property prices occurred, the economy was also struggling under the lagged effect of high real interest rates.

A third difference between the two cycles is in the underlying inflationary environment. Two aspects are important here: the inflation rate preceding the boom and the inflation rate after the boom. First, as discussed above, by the mid 1980s, Australia had already experienced over a decade of high inflation, and leveraged asset purchases appeared to many to be a successful investment strategy. This perception helped propel the boom. Second, while inflation was higher going into the 1980s boom than it was going into the 1970s boom, it was much lower during the bust; over the period 1974 to 1977, the inflation rate averaged 14%, while over the period 1990 to 1993, it averaged just 3%. So while the fall in real commercial property prices was of a similar magnitude in both cases (51% compared to 65%; see Table 2), the percentage fall in the nominal asset price was almost three times as large after the 1980s boom (21% in the 1970s downturn compared to 62% in the 1990s downturn). This large fall in nominal prices contributed to the losses recorded by some financial institutions.

Figure 6

Real GDP

Cyclical peak = 100



Finally, output took much longer to recover from the recessions in 1990/91 than from the recessions in the early 1970s and early 1980s (see Figure 6). A plausible explanation is that the substantial decline in nominal property prices led to a more protracted recession. We now turn to an econometric evaluation of this explanation that allows us to control for other factors, such as monetary policy and foreign output.

3. Asset prices, real GDP growth and inflation

Our starting point is the basic model of real GDP growth originally developed by Gruen and Shuetrim (1994) and most recently updated by Gruen, Romalis and Chandra (1997). The model we use has a long-run relationship between the level of Australian non-farm real GDP and US real GDP. The dynamics of the business cycle are influenced by the growth rate of GDP in the United States, the level of real short-term interest rates in Australia, and the output of the farm sector.[16] While it would be preferable to estimate the model from the early 1970s to include the first property-price boom, earlier work suggests that the deregulation of the Australian economy and financial markets changed some of the estimated relationships. Accordingly, we estimate the model over the

[16] See Debelle and Preston (1995), de Brouwer and Romalis (1996), de Roos and Russell (1996), and Kortian and O'Regan (1996) for a detailed discussion of these relationships

period from September 1980 to March 1997. Precise details of the data series and their construction are included in the Appendix.

Model 1 in Table 3 shows the estimated parsimonious version of this model for real non-farm GDP. The coefficient on the lagged level of Australian non-farm GDP is significant, which provides evidence of cointegration between the levels of Australian and US GDP. The sum of coefficients on lags of real cash rates is significant and negative, suggesting that tight monetary policy does slow economic growth.

Table 3

Australian non-farm GDP growth regression

Variables	**Model 1**	**Model 2**	**Model 3**	**Model 4**	**Model 5**
Constant	-2.52	-2.80	-2.21	-2.58	-2.51
	(-6.98)	(-11.32)	(-9.88)	(-10.50)	(-6.99)
Lagged Australian GDP	-0.41	-0.46	-0.36	-0.42	-0.41
(log level)	(-6.90)	(-11.22)	(-10.09)	(-10.48)	(-6.95)
Lagged US GDP	0.50	0.56	0.44	0.51	0.50
(log level)	(6.97)	(11.32)	(10.01)	(10.53)	(7.00)
Real cash rate	-0.33	-0.34	-0.29	-0.30	-0.32
(lags 2 to 6)	{0.00}	{0.00}	{0.00}	{0.00}	{0.00}
Farm output % change	0.005	0.01	0.005	0.01	0.01
(lags 2 to 4)	{0.03}	{0.00}	{0.00}	{0.00}	{0.01}
US GDP % change	-0.46	-0.65	-0.38	-0.51	-0.55
(lags 0 to 4)	{0.00}	{0.00}	{0.00}	{0.00}	{0.00}
Variable for cumulative falls in nominal commercial property prices		-0.01 (-4.04)			
Real asset prices (log change)					
Equity (lags 2 to 8)			0.04 (2.16) {0.00}		
Commercial property (lags 0 to 2)				0.04 (3.07) {0.00}	
Dwellings (lags 3 to 5)					-0.04 (-1.16) {0.01}
$\bar{R}^2$	0.63	0.66	0.64	0.67	0.69
LM test for 4th-order autocorrelation	6.76 {0.15}	10.74 {0.03}	12.26 {0.02}	8.75 {0.07}	6.00 {0.20}

Notes:

(a) Models estimated using OLS on quarterly data over the period 1980:3 to 1997:1. Numbers in parentheses () are t-statistics. Numbers in braces { } are p-values of the joint significance test that all lags of the variable are equal to zero.

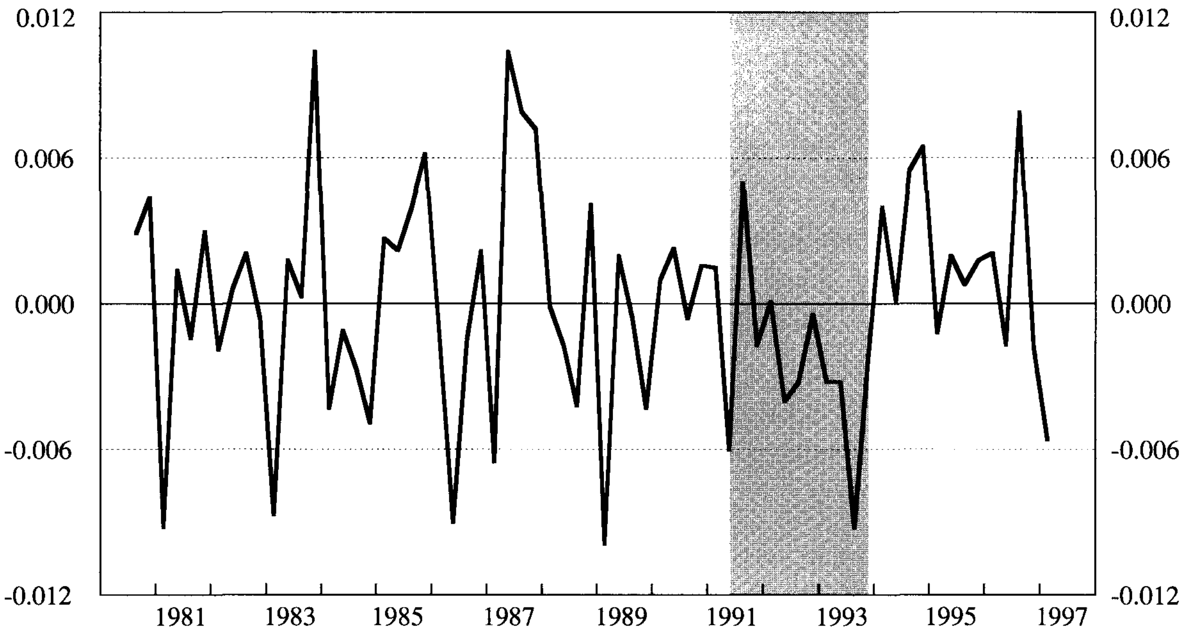
(b) Standard errors obtained by using the ROBUSTERRORS command in RATS when there was evidence of 4th-order autocorrelation in the residuals at the 10% significance level.

The residuals from Model 1 are plotted in Figure 7. While the residuals exhibit no evidence of serial correlation, the residuals from the period from June 1991 to December 1993 are especially noticeable. During this period (shown as the shaded portion of Figure 7), the model *over-predicts* economic growth in 9 of the 11 quarters.[17] The discussion in Sections 1 and 2 suggests that this unlikely outcome is the result of falling property prices, and more particularly, falling nominal property prices.

To test this proposition we added a variable to the model that takes a value of zero when the nominal price of commercial property is either rising or steady. Otherwise, the variable takes the value of the percentage fall from the previous peak in the nominal price when prices have been falling consistently from that peak. This variable captures the special effect of falling nominal property prices discussed in Sections 1 and 2 and represented by the second element on the right-hand-side of Equation (1). The results are reported as model 2 in Table 3. The coefficient on this variable is negative and significant; indicating that falls in property prices slow economic growth. The size of the coefficient is also economically quite important. By December 1992 the nominal price had fallen by almost 60% from its peak in September 1989, which means that quarterly growth was 0.6 of a percentage point below what it otherwise would have been.[18]

Figure 7

Residuals from the benchmark parsimonious model



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[17] If we include a dummy variable for the period, it has a negative sign and is significant at the 1% level.

[18] We examined a model which used a standard dummy variable in place of our preferred measure of nominal price falls; that is a dummy variable that takes the value of one when the nominal price of commercial property is falling, and zero otherwise. The results imply that a fall in nominal prices in any quarter reduces quarterly growth by 0.4 of a percentage point. When this dummy variable was added to model 2 it was not significant, while our preferred measure of nominal price falls remained significant. In support of our specification we also investigated models which included other types of dummy variables. A standard dummy variable for nominal price falls interacted with quarterly nominal price changes was not significant. A standard dummy variable for real price falls was not significant. We examined a version of model 2 which also included a variable that measured cumulative property-price increases since previous troughs; while the coefficient on this additional variable was significant and negative it was small in absolute value.

We take this result, together with the pattern of the above residuals, as evidence that the property-price cycle had a significant impact on the business cycle. To a considerable degree the protracted nature of the early-1990 recession can be attributed to the working out of the balance-sheet problems generated by the decline in nominal property prices and the reduction in financial intermediation in the first half of the 1990s.

To further examine the links between asset-price changes and GDP growth, we estimate variants of the benchmark model by including, one at a time, changes in the various real asset-price indices (models 3 through 5). The discussion earlier in the paper suggested that growth might be positively related to growth in real asset prices in the short run. To test this hypothesis we supplement the basic growth model with up to eight lags of the log difference of each real asset-price series. We report the results for the parsimonious regressions.

As models 3 and 4 show, both equity prices and commercial property prices are significant and the sum of the coefficients is positive. For equity prices, coefficients on only the second and eighth lags are individually significant, whereas for commercial property all the lags in the parsimonious model are individually significant. The sum of coefficients on lags of dwelling prices is negative but insignificantly different from zero, although individual coefficients are all significant. In each of the three cases, the inclusion of asset prices increases the goodness of fit of the model as measured by the adjusted R-squared.

It is important to note that we are measuring the combined effect of two different types of relationships between asset prices and growth. The first relationship is the direct effect of changes in asset prices on growth through the wealth effect and the cost of capital. The second relationship between asset prices and growth is not causal; instead, it reflects the fact that asset markets are forward looking. Changes in real asset prices should reflect additional information (which is not available to the econometrician) about future growth prospects. This "informational role" should be more relevant for equity markets because they are more liquid and cover a wider range of economic activities.

When we compare the dynamic structure of models 3 and 4 we note that only the earlier lags on the coefficients of commercial property prices are significant and these coefficients are larger than those on early lags of equity prices. Whereas longer lags of changes in property prices are not significant, the coefficient on the eighth lagged change in real equity prices is positive and individually significant (at the 10% level). A plausible explanation of these results is that the direct effects of commercial-property prices are relatively strong and occur rapidly. Also, the ability of asset prices to provide *extra* information on future growth rates is likely to be evident over longer horizons, as appears to be the case for equity prices.

In summary, equity prices appear to have informational value in predicting future growth in the economy. More importantly, the above results support the idea that the working out of the asset-price bubble of the late 1980s helps explain why economic growth in the early 1990s was slower than that suggested by standard explanations of the pace of economic growth.

As part of our empirical work, we also added percentage changes in the real asset-price series to models of goods and services price inflation. These models include lagged levels and changes in unit labour costs and import prices as well as the output gap. In none of the models that we estimated did we find a significant role for any of the asset-price series. This is hardly surprising, since asset prices are expected to have their main effect on inflation through affecting the dynamics of the business cycle. Falling asset prices which cause a protracted recession are likely to lead to a decline in wage pressures and firms' margins, and thus to lower inflation. However, once we have controlled for these effects, there is little, if any, independent role left for asset prices.

4. Lessons for monetary policy

We draw four broad lessons from the above discussion:

1. Property-price cycles and credit cycles go hand in hand.

Over the past thirty years, the two large cycles in the ratio of credit to GDP have coincided with the two large cycles in commercial-property prices. Furthermore, downturns in property prices have been associated with slow recovery phases in activity. These links make property prices particularly important to monetary policy.

2. Monetary policy can burst property-price bubbles. Under some circumstances, it may make sense to do so, even if it means that expected inflation is below the central bank's target. The case for attempting to influence equity prices is weaker than for property prices.

Monetary policy can burst bubbles in property prices by increasing the cost of speculative behaviour and by slowing down the growth rate of economic activity. If the only instrument of policy is short-term interest rates, a protracted period of tight policy is probably required. This means that when the bubble bursts, losses are likely in the financial system and economic growth is likely to be below trend. But once a bubble has emerged, avoiding such an outcome will prove even more difficult. By bringing forward the collapse of the bubble, monetary policy can reduce the scale of the inevitable slowdown in economic activity. The main difficulties with such a policy are in identifying that a bubble does indeed exist and generating the necessary public acceptance of a period of tight monetary policy.

In contrast, bubbles in equity markets are likely to have much weaker implications for the health of the financial system. While policy-makers need to be alert to changes in patterns of financial intermediation which might change this situation, Australian experience suggests that equity-price bubbles can burst without detrimental effects on financial intermediation.

3. Financial liberalisation matters (but not as much as we might think).

The liberalisation of the financial sector played a role in the property boom of the late 1980s. Competition, combined with a lack of credit-assessment skills, saw a lowering of credit standards and this contributed to a larger increase in the ratio of credit to GDP than would likely have occurred in a more regulated environment. However, it is easy to overstate the importance of financial deregulation. The property boom of the 1970s shared many of the characteristics of the boom of the late 1980s. Regulation prevented the banks from being full participants in the earlier boom, but this simply encouraged other institutions to supply the credit. Despite the regulation there was still a very large cycle in credit and the economy experienced the contractionary effects of the bust of the property cycle and reduced financial intermediation.

One, often overlooked, aspect of financial deregulation that can contribute to booms in property prices is the effect that deregulation has on the underlying value of commercial property. The limited deregulation of the early 1970s and the comprehensive deregulation of the mid-1980s led to a significant increase in the demand for prime CBD office space. This put upward pressure on the fundamental price, which made it easier for a bubble to emerge.

4. The inflationary environment matters for bubbles.

Low inflation should make bubbles less likely but if they emerge nevertheless, low inflation can make their bursting more costly. If an asset-price bubble does occur, a low-inflation environment makes it more likely that the inevitable correction in *real* prices will occur through a decline in *nominal* prices. This increases the risk of financial-system instability and can lead to a protracted period of weak growth in financial intermediation. This increases the case for the central bank to tighten policy relatively early in the life of a property-price bubble.

Appendix : The data

All data are quarterly except those marked with a * which are measured for the year ending 30th June.

Equity price index. Quarterly average of the Australian All Ordinaries share price index. Source: Datastream code – AUSTALL.

Capital Value Indicator (CVI) for Sydney CBD property. The average capital value per square metre of net lettable area. Source: Jones Lang Wootton (JLW) Advisory Services.

Dwelling Price Index. Constructed by splicing together the most appropriate prevailing house price series. Sources: BIS Shrapnel average house prices for Sydney and Melbourne (March 1960 to December 1969); Abelson weighted average of house prices (March 1970 to September 1981); REIA median prices of established houses (December 1981 to March 1984); Commonwealth Bank of Australia Housing Industry Association (CBA-HIA) (April 1984 to present).

Median prices of established houses in Sydney. Source: Real Estate Institute of Australia (REIA).

Real asset prices series are calculated by deflating nominal prices by the implicit price deflator derived from the seasonally adjusted expenditure measure of GDP. Source: ABS National Accounts Cat. No. 5206.0.

Real gross domestic product, GDP(E). Expenditure measure of GDP (seasonally adjusted). Source: ABS National Accounts Cat. No. 5206.0.

Credit to the private sector by all financial intermediaries. Loans and advances, plus bank bills discounted (break-adjusted, seasonally adjusted). Prior to August 1976, annual seasonally adjusted data is interpolated. Source: Reserve Bank of Australia Bulletin - Table D3.

Total assets of financial intermediaries. The sum of total assets of banks and all other financial institutions. These series are taken as a ratio to the income measure of GDP. Source: Reserve Bank of Australia Bulletin - Table D5.

*Average return on shareholder's funds (major banks)**. Source: Australian Stock Exchange.

*Prime lending rate**. Business indicator rate on bank's large, variable rate business loans. Source: Reserve Bank of Australia Bulletin - Table F4.

Output Gap. Deviation of GDP(A) from potential output based on linked peaks of multifactor productivity. Source: Reserve Bank of Australia.

Real gross domestic product, GDP(A). Average of output, production and income measures of GDP (seasonally adjusted). This series combines farm and non-farm output measures used in regression analysis. Source: ABS National Accounts Cat. No. 5206.0.

Underlying Consumer Price Index (CPI). Treasury underlying consumer price index. Source: Consumer Price Index, ABS Cat. No. 6410.

Cash rate. Source: Reserve Bank of Australia Bulletin - Table F1.

United States real GDP. US GDP chain linked. Source: Datastream code - USGDP...D.

Stock price index for Australian property developers. Source: Datastream code - PRPTYAU.

Estimating Sydney commercial property prices between 1973 and 1977

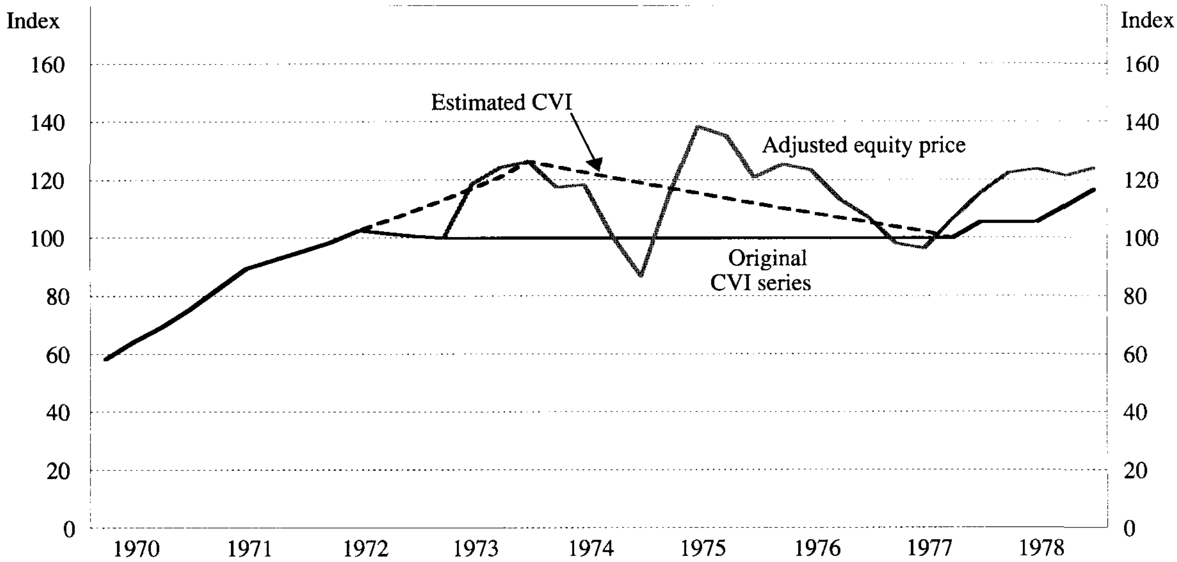
Figure A1 shows the original series for the Sydney commercial property Capital Value Index. Between March 1973 and September 1977 the original series exhibits no variation because only one valuation was conducted during this period. Furthermore, the small nominal price fall from June 1972 to March 1973 is inconsistent with the detailed discussion in Daly (1982). We replace the original series with an estimate from June 1972 to September 1977. In line with Daly (1982) our estimate peaks in December 1973 and then declines linearly back to the level of the original series in

September 1977. The level of the peak is obtained by setting the growth rate of the estimated CVI equal to the average growth rate of the original series over the year 1971/72.

Figure A1

Commercial property prices – original and estimated CVI

March 1973 = 100



For comparison we also show the movement in an index of equity prices of companies that were involved in the development of commercial property in Sydney (available from March 1973). We have adjusted this series to account for its correlation with the aggregate Australian stock price index (by taking the residuals from a regression of the commercial property stock price index on the All Ordinaries index). Compared to our estimated CVI, the adjusted property equity price index displays more volatility, but a similar growth rate before and after the peak in December 1973.

Timing of asset price cycles

	Equity	Commercial property	Dwellings
		Starting dates of each phase	
Early-1970s cycle			
(in between)			66:3
Upturn	66:4	68:1	71:3
Downturn	70:1	73:4	74:2
(in between)	74:4	77:3	
Late-1970s/Early-1980s cycle			
Upturn	78:1	78:2	79:3
Downturn	81:2	82:1	81:4
(in between)	82:3	85:1	84:4
Late-1980s cycle			
Upturn	84:3	87:1	87:2
Downturn	87:3	89:3	89:1
(in between)	92:4	93:3	93:1

Note: The cycles are delineated according to either turning points or points of inflection in nominal asset prices. If instead we used real asset prices the timing would change marginally for only a few phases.

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Asset inflation in the Netherlands: assessment, economic risks and monetary policy implications

Jeannette Capel and Aerdt Houben*

Introduction

In recent times, equity and house prices in the Netherlands have soared. By late 1997 the AEX index of the Amsterdam Exchanges hovered around the 900 mark, having started the year at 634, reaching a maximum of 1011 on 7th August. House prices went up some 10% in 1996, a rate of increase that continued in the first half of 1997. This has led many to wonder whether in the early months of 1997 the Netherlands was subject to asset inflation. Asset inflation (or an asset bubble) occurs when the prices of financial assets rise above their underlying or intrinsic value.

This article addresses that question and discusses the economic risks and monetary policy consequences of asset inflation. Section 1 looks at the fundamentals that determine the intrinsic value of equities and houses. On the basis of this analytical approach, Section 2 examines whether the recent price rises in the Dutch financial markets are inflationary in nature or whether they can be attributed to improved fundamentals. Subsequently, Section 3 describes the economic risks of asset inflation for the Netherlands. As was shown dramatically by the 1929 Wall Street crash and the ensuing global depression, a bursting asset bubble may prompt a financial crisis and a collapse in economic activity. There is now consensus that central banks can mitigate the adverse effects of a bursting bubble by easing monetary policy. A more intriguing policy question, which is addressed in Section 4, is whether a central bank should wait until the bubble has actually burst, or whether it should play a more active role in preventing bubbles from arising in the first place.

1. The intrinsic value of equity and houses

1.1 Definition of asset inflation

Asset inflation occurs when the prices of financial assets (represented by P_t) are rising even though they are already above their intrinsic or underlying value (V_t):[1]

$$dP_t / dt > 0 \text{ and } P_t > V_t \tag{1}$$

Hence, to establish asset inflation, the intrinsic value of equities and houses must be first determined.

1.2 Equities

In finance theory, the intrinsic value of equity (V_t) is determined on the basis of the net present value of total expected future income. The calculation is usually based on the Gordon model,

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1 If $dP_t / dt > 0$ but $P_t < V_t$, equity or house prices are on the increase but moving towards equilibrium.

which assumes that dividend payments D increase by a constant growth rate g from time t, and that the equity is held for an infinite period of time.[2]

It then follows that:

$$V_t = \frac{D_t}{r - g} \tag{2}$$

where r is the required stock market return.[3] Equation (2) applies to individual stocks as well as to equity portfolios, such as the portfolio of 25 active equities underlying the Dutch AEX stock market index.

The Capital Asset Pricing Model (CAPM) shows that the required stock market return r is the sum of the risk-free interest rate (the minimum remuneration for making capital available) and a risk premium. The latter depends on the degree of market risk to which the equity or equity portfolio is subject (the so-called beta) as well as the degree of risk aversion (which determines the required remuneration per risk unit).[4]

Accordingly, the intrinsic value of an equity or of an equity portfolio is determined in part by three non-observable variables (viz. growth expectations, the market risk involved in the equity, and the degree of risk aversion), which are known to vary over time. Historical data may thus give a wrong impression of the current situation, making it impossible to pinpoint the exact intrinsic value of an equity or an equity portfolio.

For the same reasons, it is also impossible to determine the precise underlying value of the price-earnings (P/E) ratio, which is often used by investors in practice. To determine the underlying P/E ratio the following assumptions are required (see French and Poterba (1991), p. 354): a firm reinvests a fraction k of its profits E and this investment achieves a supernormal return of r^* (that is one which exceeds r, the required stock market return) during period T. The firm's profits then increase by rate g, which equals kr^*. The remaining profits of $(1\text{-}k)E$ are paid out as dividends. The intrinsic value of the P/E-ratio is then

$$P/E = [1+kT(r^* - r)] \,/\, r = [1+T(g-kr)] \,/\, r \tag{3}$$

which again depends on the non-observable values of the expected growth of profits g and the required stock market return r.[5] In addition, P/E ratios may differ considerably between countries and industries.[6] For these reasons, it is impossible to ascertain unequivocally whether increases in certain

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2 Obviously, as the Gordon model is a simplification of reality, it is used here solely to indicate that growth expectations and the rate of return required by market participants determine the intrinsic value of equities. More realistic models would call for a more complex presentation, without providing substantial new insights.

3 The growth rate g and required market return r should both be either in nominal or in real (i.e., inflation-adjusted) terms.

4 The Arbitrage Pricing Theory (APT) gives a similar definition of required market return. Unlike the CAPM model's abstract concept of market risk, the APT model distinguishes between separate risk factors, such as disappointing growth and inflation, so that r is the sum of the risk-free interest rate and various risk premia (relating to the various risk factors). These risk premia are ultimately determined, as in the CAPM model, by the degree of risk entailed in the equity and the extent of risk aversion.

5 If an infinite time horizon is assumed, $P/E = (1\text{-}k) \,/\, (r\text{-}kr^*) = (1\text{-}k) \,/\, (r\text{-}g)$, in accordance with equation (2). It appears, furthermore, that if no profits are retained (i.e. if $k = 0$), $P/E = 1/r$, then $E/P = D/P = r$; in other words, the dividend equals the required stock market return if all profits are disbursed.

6 For example, French and Poterba (1991) show that, by comparison with the United States, Japan has high P/E ratios which are partly accounted for by differences in accounting methods and tax rules.

P/E ratios reflect improved fundamentals or asset inflation.

Despite these shortcomings, equation (3) does provide a rough method for ascertaining asset inflation. On the basis of the *actual* P/E ratio and a reasonable assumption regarding the required stock market return, the *implicit* growth rate of profits can be calculated. It is then possible to gauge whether this implicit growth rate is realistic, given the cyclical situation.[7] Section 2 applies this method to the Netherlands.

1.3 Real estate

Real estate may be seen as a financial asset, but most house-owners consider it a (durable) consumer good. Unlike other markets for consumer durables, however, the housing market is characterised by an almost completely inelastic short-term supply curve, which implies that virtually every change in demand will lead to a change in prices. Demand changes may stem from real economic or monetary factors. If house prices are pushed above their intrinsic value by monetary factors such as overly generous mortgage lending, one speaks of asset inflation.

The intrinsic value of houses in the Netherlands may have risen due to the following real causes: (1) weak supply because housebuilding has not kept pace with the increasing number of households; (2) an increasing quality of houses; and (3) growing demand for home ownership because it has become cheaper to own a home than to rent one. Section 2 will examine whether these real factors can account for the recent house price rises in the Netherlands or whether there is evidence of asset inflation.[8]

2. Is the Netherlands subject to asset inflation?

2.1 Introduction

Following the above discussion of the fundamental determinants of the intrinsic value of equity and houses, this section deals with the question whether the Netherlands is currently facing asset inflation. The next section assesses whether the recent rises in the P/E ratios of equities can reasonably be attributed to a lower required stock market return and/or a higher expected profit growth. If that is not the case, there is an indication of asset inflation. Movements in house prices are then reviewed in the light of developments in (1) the supply of houses, (2) the quality of owner-occupied dwellings and (3) the cost of home-ownership as compared to rents.

2.2 Equities

Dutch equity market developments

Figure 1 shows that the stock exchange boom until last August was an international phenomenon, but that the Dutch index has recently gone up much more markedly than those of the United States, Japan, the United Kingdom and Germany. The relatively low level of (risk-free) capital

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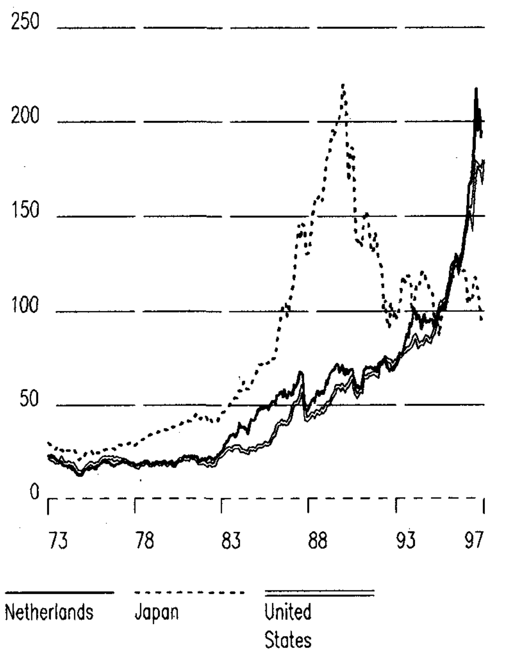
7 Historical experience indicates that P/E ratios usually peak during the early stages of cyclical upturn because of the potential for above-average growth in the medium term (see, for instance, BIS (1997), p.72).

8 If a house is considered an investment object, its intrinsic value is also determined by rents and mortgage rates. The analysis of asset inflation therefore does not disregard essential information when houses are viewed solely as consumer durables.

Figure 1

International stock market indices
1995 = 100, monthly figures



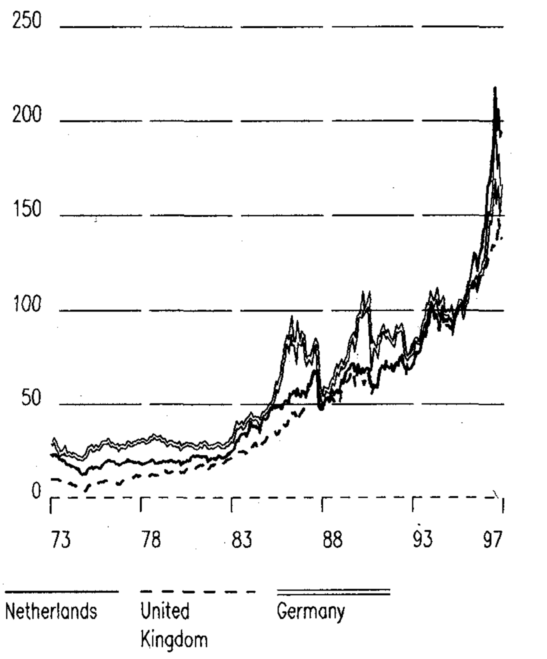


Figure 2

Long-term interest rate and stock market index
Monthly figures

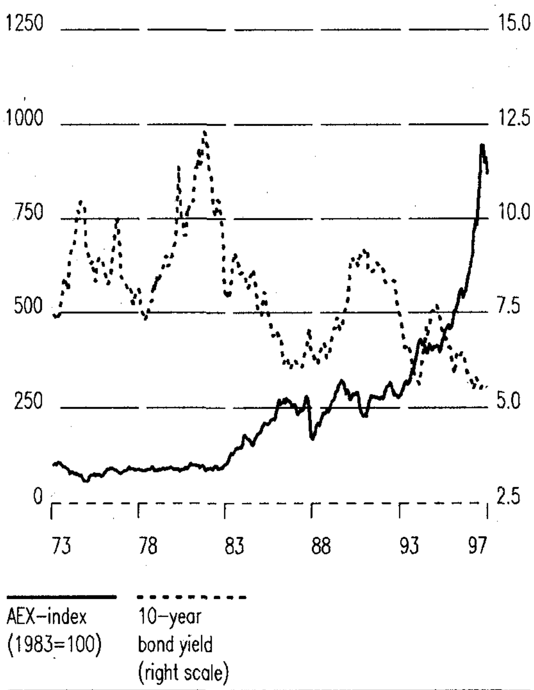
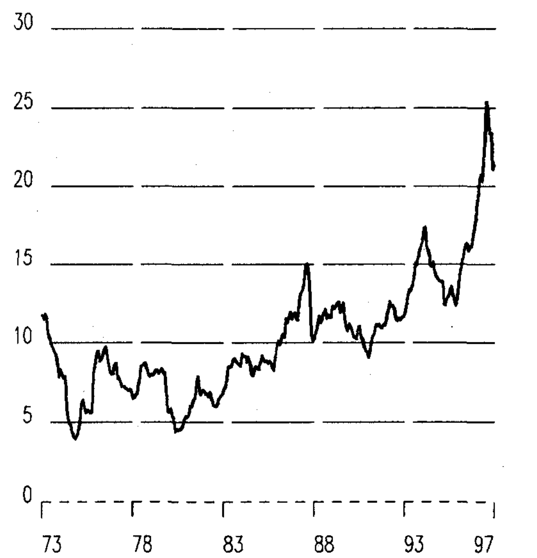


Figure 3

Price/earnings ratio of the Dutch stock market
Monthly figures



Source: For Figures 1 and 3, Datastream.

market rates seems to explain part of the Dutch equity price rises (Figure 2). In the analysis below, we will focus on P/E ratios to see whether a combination of higher expected profit growth and lower risk premium can complete the explanation. Figure 3 shows that the P/E ratio has been undergoing a trend rise since the mid-1970s which has been considerably faster in 1996 and 1997 than in previous years. For that reason we will focus on the developments in the P/E ratio since 1995.

Assumptions made and data used in the analysis

An implicit expected growth rate of profits can be calculated by using equation (3) and the realised P/E ratio and by making a reasonable assumption on the required market return r. An assessment is then made as to whether this calculated implicit rate is realistic by comparing it to a benchmark of 7% annual real profit growth. This profit growth would be realised in the Netherlands in the next ten years if GDP expands at a rate of 3% (2.7% over the past ten years) and if the profit ratio gradually increases from 8.7% (realised in 1996) to 12.4% (the maximum in the past decade).[9] This profit benchmark is exceptionally high since the Dutch economy is currently in a mature phase of the business cycle. We have deliberately chosen such a high benchmark: if the actual expected profit growth should be even higher to explain the increase in the P/E ratio, we have quite a clear indication of asset inflation.

Our calculations assume that supernormal returns can be realised during ten years (so $T = 10$, as in French and Poterba (1991)). The reinvestment rate was calculated using actual figures on the Dutch P/E ratio and the dividend/price (D/P) ratio, viz. $k = [1\text{-}P/E * D/P]$. Our reasonable assumption for the real required market return was the interest rate on the latest ten-year central government loan minus an expected inflation rate of 2%[10] plus a risk premium of 6.5% (see De Haan (1997) and Poll (1996)). However, as equities could be a more customary form of savings today than in the past, risk premia below 6.5% were also looked at. Table 1 summarises the data used.

Table 1

Data used in the calculations

	1995	**1996**	**1997**	
			7th August (max.)	18th Dec. (last)
P/E ratio	13.0	18.6	26.4	21.3
Reinvestment rate	59.3%	52.0%	52.2%	55.5%
Real risk-free interest rate	4.9%	4.2%	3.6%	3.3%

Results

The calculated implicit expected profit growth figures are presented in Table 2. If the required risk premium in the period 1995-97 had remained at its average past value of 6.5%, the expected profit growth would have needed to be between 11.6% (1995) and 22% (1997) to justify the actual values of the P/E ratio. Though such growth figures could conceivably be realised in the course of a single year, there is no question of them being realised for a period of ten years. So the next step

9 The profit ratio would then rise by an annual 4%, averaging 10.5%. The expected profit growth would then be: 1.04 *1.03 (GDP growth) - 1 ≅ 7%. As the profit ratio has averaged 9.4% over the past ten years, this estimated profit growth rate is rather high, owing to the assumptions on both GDP growth and the profit ratio.

10 2% has been the average inflation rate in the Netherlands in the past decade.

Table 2

Implicit expected profit growth figures, in percentages

Risk premium	1995	1996	1997 7th August (max.)	1997 18th Dec. (last)
6.5%	11.6	15.4	22.0	16.2
6%	10.6	14.2	20.4	14.8
5%	8.7	11.8	17.2	12.2
4%	6.9	9.4	14.1	9.5
3%	5.0	7.0	10.9	6.8
2%	3.0	4.6	7.8	4.1
1%	1.2	2.3	4.6	1.4
0%	-0.7	-0.1	1.4	-1.3

is to explore whether the high P/E ratios could be caused by a lower risk premium. From Table 2 it appears that the P/E ratio in 1995 and 1996 can only be explained if the risk premium on equity had fallen to about 4% and 3%, respectively. At the peak of the stock market on 7th August 1997, the risk premium had to be below 2% to account for the P/E ratio at the time. As it seems unlikely that the risk premium on equity had fallen so much in such a short period of time, these figures give a clear indication that the Dutch stock market was characterised by asset inflation in the summer of 1997. Since then, the P/E ratio has moved closer to its intrinsic value. Nevertheless, the current P/E ratio is still high, requiring a risk premium of about 3%, which is half of its past value.

2.3 Real estate

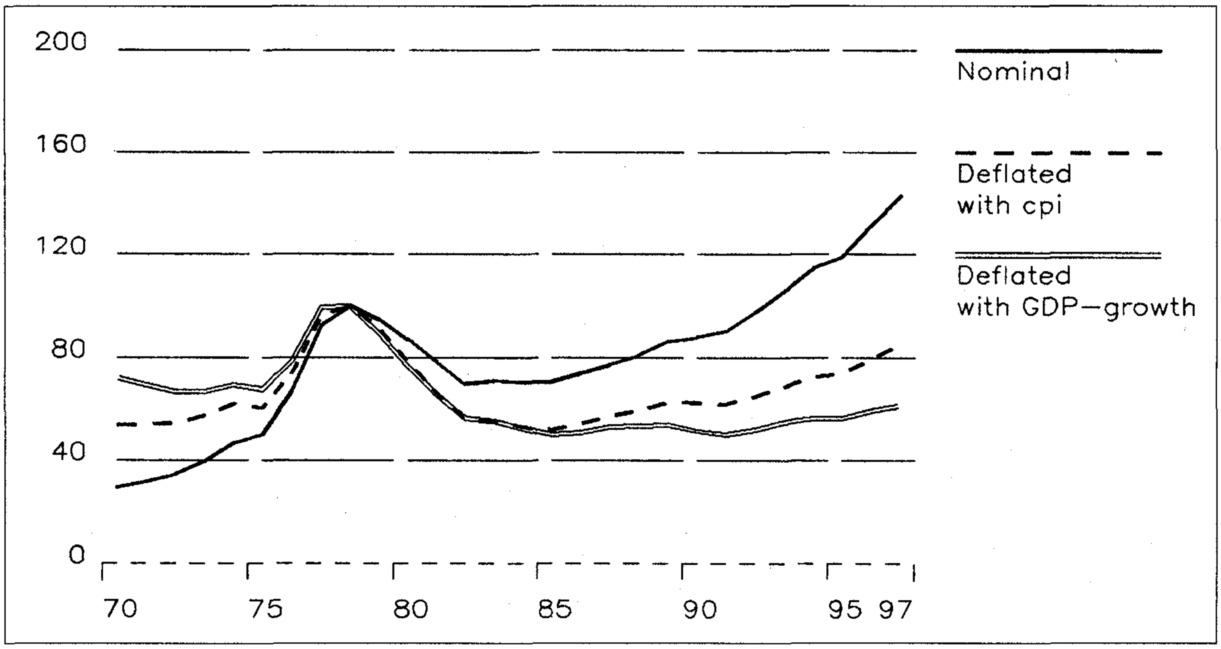
House prices in the Netherlands are currently far above the level of 1978, when the market peaked and subsequently a slump set in (Figure 4). However, Figure 4 also shows that *real* house prices, i.e. house prices deflated by the consumer price index, are still well below the 1978 peak. Considering developments in the relevant real economic factors (Section 1), it seems that the intrinsic value of houses has increased, leaving little – if any – indication for asset inflation in the housing market at present. First, there are demographic developments: the number of households has increased steadily, whereas political problems caused a slowdown in the supply of new houses. Second, new houses have become more luxurious and the quality of existing houses has been improved (e.g. dormer windows have been added or kitchens/bathrooms modernised) with the increase in prosperity. Figure 4 shows that if the nominal house price is deflated by the nominal GDP growth rate approximating the increase in prosperity and improved housing quality, house prices have barely gone up in recent years. Finally, the underlying value of houses for owner occupation should has risen because the costs of home ownership have fallen vis-à-vis the costs of renting a house. This is shown in Figure 5.[11] Given that the house price rise has been quite moderate in real terms and that there are, at least, three factors providing a justification for a higher real house price, there is little

[11] The costs-of-home-ownership index is calculated as follows: the average mortgage burden is put at 50% of the mortgage rate (after taxation) * the average house price. To this, the fiscal costs of notional rent income are added: 60% of the average house price * the rate of notional rent income (which varied between 1.5% and 2% in recent years) * 50% tax (i.e. the marginal tax rate to which most home owners are subject). Figure 5 compares this to the rental cost index, published by Statistics Netherlands.

evidence of asset inflation in the Dutch real estate market. This is in contrast with our conclusion above that the Dutch stock market was – and probably still is – subject to asset inflation.

Figure 4

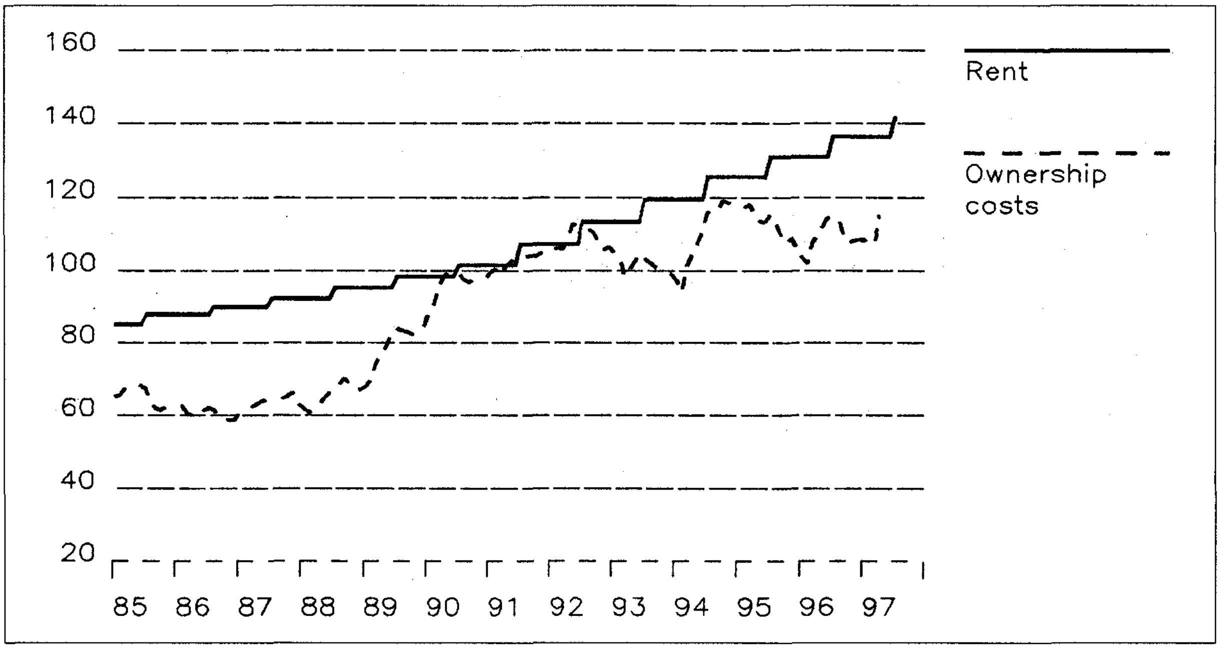
House prices relative to economic growth and consumer prices
1978 = 100



Sources: For 1970-75, CPB estimates, NVM thereafter.

Figure 5

Home ownership costs versus rental costs
1990 = 100



3. When the bubble bursts: the risks of asset inflation in the Netherlands

3.1 Theory: the monetarist and the financial stability approaches

The main danger of asset inflation is that it may generate a future equity price decline or a fall in house prices, thereby dampening real economic activity. The economic literature provides two main theories on this: the monetarist theory (based on Friedman and Schwartz (1963)) and the financial fragility theory (put forward by Minsky (1977)). According to the monetarists, an asset bubble may arise even though the risk of a crash has been correctly priced ex ante by rational economic agents. While the bubble lasts, the average return is higher than the risk-free interest rate because the chances of a crash have been correctly discounted in a risk premium. When the bubble bursts, a financial crisis may – but does not necessarily – occur. If economic agents continue to have confidence in the liquidity of their bank deposits, there is only a "pseudo" financial crisis (Schwartz (1987)). Private expenditure will be lower because the net capital of households and firms has decreased, but there is no danger of a money supply contraction and therefore no need for the monetary authorities to intervene. A "real" financial crisis occurs if a stock exchange crash causes the public to lose faith in the banking sector prompting a massive withdrawal of deposits. If this leads to bank failures and a contraction of the money supply, the central bank should act as lender of last resort to avoid deflation.[12] In the monetarist view, it is unlikely that a financial crisis will cause an economic downturn as long as monetary policy is adequate, although such a crisis may reinforce and prolong a downturn.[13] Indeed, monetarists argue that the causality is often the other way around: cyclical weakening sets in motion a confidence crisis which may cause problems for banks. For these reasons, monetarists do not see a necessary link between asset bubbles and business cycles.

Most proponents of the financial fragility approach see asset bubbles as the product of an irrational mania on the part of investors, so that risk is being under-priced ex ante. Financial fragility is easily built up during boom periods when rising equity and house prices, as well as inflation, stimulate excessive lending by banks, i.e. lending which exceeds the expected income flows of households and businesses (this is known in the literature as speculative or Ponzi financing). Under these conditions, it is plausible that a stock or real estate market crash leads to a banking crisis. Consequently, consumer spending and investment, as well as output and employment, are depressed because of the decline in private agents' net capital and the reduction in lending stemming from the banking crisis.[14] In short, contrary to the monetarist approach, the financial fragility approach perceives a clearly discernible connection between asset bubbles and business cycles.

Banking crises play a major role in both the monetarist and the financial fragility approach. The next section addresses the sensitivity of Dutch banks to a stock exchange crash or a real estate crisis. In addition, both approaches emphasise that a fall in house or equity prices impairs the net capital position of private agents, which in turn adversely affects private consumption and business investment. Section 3.3 presents empirical evidence on the size of these effects in the Netherlands.

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[12] According to monetarists, financial instability can be largely prevented by pursuing price stability. If inflation is volatile, so are real interest rates, making it difficult for banks and other financial institutions to properly assess the creditworthiness of their debtors, and increasing the risk of bad debts.

[13] Bernanke (1983) shows that the length of the depression in the 1930s was due to the higher costs of financial intermediation, which prevented households, farmers and small businesses from obtaining credit and forced them to cut down on their spending.

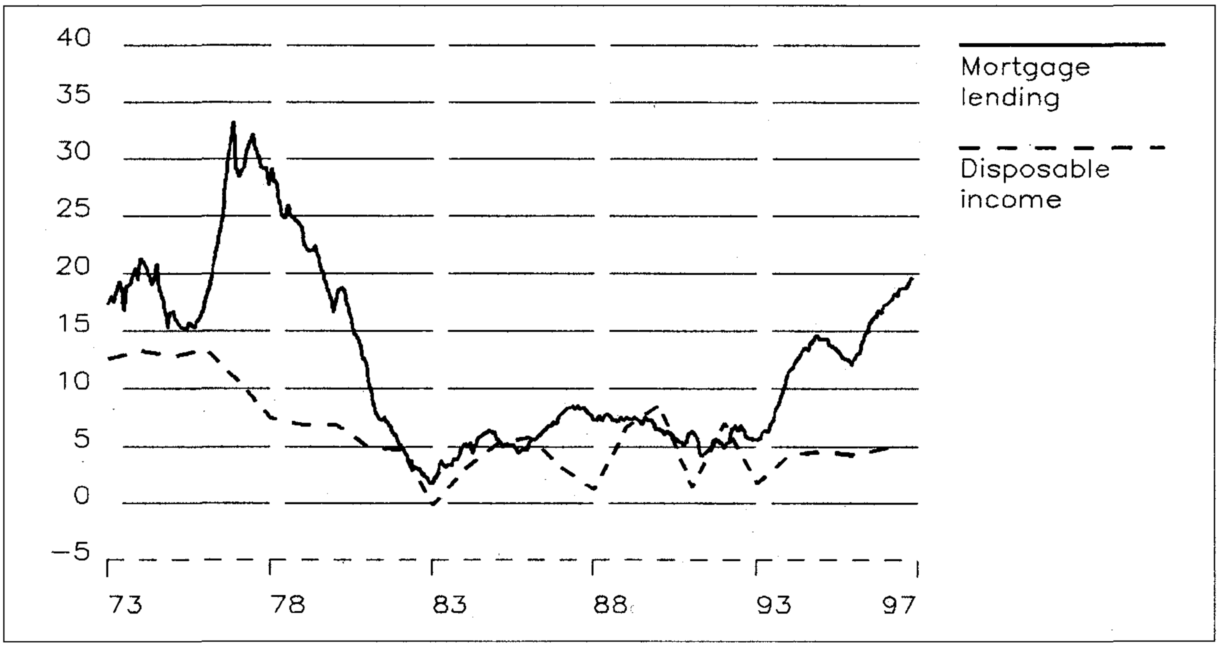
[14] This process may be reinforced by debt deflation (see Fisher (1933)): when debtors repay their debts, the money supply and the price level go down, generating an increase in the *real* value of private agents' debts.

3.2 Banking crises

A bursting asset bubble is likely to depress the banking sector balance sheet for two reasons. First, an asset crash has a *direct* impact on the balance sheet if banks invest in equity or real estate. In addition, there may be an *indirect* effect if during the preceding boom period banks have engaged in speculative or Ponzi financing; i.e. have granted loans which form an excessive burden on the expected future income and cash flows of households and businesses, on the assumption that the collateral underlying the individual loans (equities in the case of securities-based lending, real estate in the case of mortgages) will retain its value or even increase in value as time goes by. If a crash takes place, such loans lose their value: the contraction of economic activity saddles households and businesses with payment problems, while at the same time the collateral – the loan's liquidation value – has decreased in value. If the banking sector balance sheet deteriorates, a confidence crisis could emerge, resulting in depositor withdrawals and the possible failure of banks. Banks with a low solvency ratio will be the first to go. All this warrants the conclusion that the risk of a crash-induced financial crisis depends on (1) the extent to which banks own equities or real estate; (2) whether banks have undertaken excessive lending as defined above; and (3) their solvency ratio.

Figure 6

Mortgage lending and disposable income
Annual percentage changes



It seems unlikely that a bursting asset bubble can set off a banking crisis in the Netherlands. Dutch banks have small equity holdings (less than 1% of the total value of outstanding equities), especially compared to banks in other countries. At the same time, the solvency ratio of Dutch banks is fairly high: around 12%, well in excess of the 8% BIS standard (BIS (1997)). No empirical data are readily available on the criteria which banks apply to their lending operations, so we cannot judge to what extent this lending can be called "excessive". It appears, however, that the banks' vulnerability to crises in the real estate market may have risen. Figure 6 shows that the recent growth in mortgage lending has far exceeded the development in disposable income. This is due to the fact that for some years now, both double (partner) and temporary incomes are taken fully into account in determining the maximum mortgage. In addition, more can be borrowed in relation to the value of collateral: in the past 70% of the forced-sale value of a house was the criterion, whereas nowadays it is generally 125% with outliers to 150%. Finally, the forced-sale value, expressed as a percentage of the purchase price, has gone up over recent years. On the other hand, the increase in vulnerability should not be exaggerated since the average house-buyer brings in more own funds today (25-30%) than in the 1970s. Moreover, some banks have recently tightened their acceptance policy by including an assessment as to whether a potential debtor can still meet his debt service obligations if interest rates go up.

3.3 Consumption

Generally speaking, rising asset prices will stimulate consumer spending because (1) households' net capital expands; (2) a buoyant stock exchange climate and rising house prices usually boost consumer confidence; and (3) credit becomes more readily available (Section 3.2). For those very same reasons, private consumption will suffer when a crisis arises in the equity or real estate market. The two crises are likely to have different quantitative effects. Equity ownership is concentrated in the better-off part of the population (whose propensity to consume out of capital is low), while home ownership is spread much more equally. This implies that a loss of capital ensuing from a real estate crisis would have a stronger impact on consumption than an equally large loss resulting from a stock exchange crash.

The effect of a crash in the equity or real estate market on consumer spending can be quantified by examining the statistical link between capital and consumption, but there are two caveats. First, where equities are concerned, a positive statistical link need not be an indication of a wealth effect. The causality may be the other way around due to the leading indicator properties of consumption. In Poterba and Samwick (1995), it was shown empirically that in the United States higher consumption was the cause, rather than the result, of higher equity prices. Second, it could well be that consumer spending reacts disproportionately strongly to asset market crashes as compared to modest price falls (i.e., wealth effects are non-linear). The reason is that the increased uncertainty makes for a greater propensity to save (Dornbusch and Fischer (1987), p. 276).

Equities

With these two considerations in mind, an empirical estimation of the effect on consumption of a 20% fall in equity prices can be derived from the MORKMON model of the Nederlandsche Bank.[15] The first year the estimated effect on the volume of Dutch private consumption is negligible, while after four years consumption will be a total of 0.3% lower. It is worth noting that the strongest influence of the crash is exercised through the slump in economic activity abroad. The direct wealth effect is very small (not more than a cumulative decline of 0.1% after four years) since equities comprise a relatively small share of Dutch households' financial assets.[16]

Houses

According to the MORKMON model, a 10% fall in house prices has a much stronger negative effect on private consumption (of -0.1% after one year to a cumulative -0.8% after four years). This total effect is made up of a direct wealth effect on consumption (-0.1% after one year; a cumulative -0.4% over the next three years) and indirect effects through higher unemployment, stronger wage moderation and hence decreased purchasing power. The direct wealth effect of a 10%

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[15] The 20% fall in equity prices worldwide was first translated into the relevant exogenous variables for the Dutch economy using the NIGEM model.

[16] Wealth effects in the Netherlands are small in general. The Nederlandsche Bank's macro-economic model for the European Union (EUROMON), shows that consumption in the Netherlands, as compared to the comsumption of several other European countries, has the lowest wealth elasticity.

fall in house prices is much stronger than that of a 20% drop in equity prices because home ownership is more equally spread over the Dutch population.

3.4 Investment

The economic literature is ambiguous as the relation between the level of equity prices and investment is unclear. According to Tobin's q theory, it is lucrative to invest if the ratio q between a business's market value and its replacement value exceeds 1. So equity price declines, which cause $q < 1$ for certain businesses, should have a negative impact on investment. However, as pointed out by Blanchard, Rhee & Summers (1993) and others, the evaluation by the manager of the marginal investment project (q^*) may deviate from the market's valuation (q). This may be due to the fact that managers are better informed than investors or that the market value is unduly high relative to the fundamentals because of rational or irrational bubbles (see Section 3.1). The cause of the discrepancy between q and q^* determines which of the two should underlie the investment decision (see Blanchard, Rhee and Summers (1993)). There is, therefore, no obvious theoretical relation between equity prices and investment.

Empirical research shows that market value and hence equity prices play no more than a limited role in investment decisions (see, for example, Blanchard, Rhee and Summers (1993) for research on the United States). Regarding Dutch businesses, Tobin's q has not been found to have any significant influence on investment behaviour (De Haan (1997), Van Ees and Garretsen (1994) and Van Els and Vlaar (1996)), since most investment is financed with internal funds (i.e., retained profits). In the period 1985-90, over 50% of the investment by Dutch businesses was financed in this way (Van Ees and Garretsen (1994)).[17] External financing is often considered too expensive because of agency problems, information asymmetry and taxation. These objections are stronger for equity financing than for bank loans.[18]

Still, in spite of the limited role played by equity in the financing of investment,[19] it is possible that a stock market crash could have an indirect negative effect on investment, in that it could impair producer confidence and further restrict access to bank loans, thereby limiting businesses in their investment opportunities. The fall in consumer spending may also be expected to cause a contraction of investment.

4. Implications of asset price inflation for monetary policy

4.1 Introduction

In the event of indications of asset price inflation, the question arises whether, and how, the central bank should react. This fundamental question can also be addressed by turning the argument around. Indeed, it is commonly accepted that, once a speculative asset price bubble bursts, central banks have a role in mitigating any adverse impact on the real economy. This is not only a key lesson drawn by monetarists, among others, from the stock market crash in 1929 (in particular with

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[17] Note that this study dealt with quoted companies. Small firms rely even more heavily on internal financing.

[18] As a consequence, internal financing is particularly important for the investment by businesses that are subject to credit constraints (Van Ees and Garretsen (1994) and De Haan (1997), the latter observing that one-third of Dutch businesses are facing credit restrictions).

[19] New equity issues play a fairly volatile role over time in the financing of investment, averaging around 10% (De Haan (1997), p. 105).

respect to the importance of avoiding a money supply collapse induced by the weakening of financial institutions), but it is also a major theme in the theories on financial instability (Minsky (1986)) and asymmetric information (Mishkin (1991)). Does this not imply that these same monetary authorities have a task in preventing the emergence of such bubbles in the first place? The following subsection addresses this question. First, the principal monetary policy objective of ensuring price stability is translated into operational terms and subjected to a broad review. Subsequently, on the basis of the analysis in previous sections, the different channels through which asset prices influence consumer prices are mapped out. As monetary policy primarily focuses on consumer prices, this provides a conceptual framework for identifying the central bank's possible role when asset price inflation occurs.

4.2 Asset price inflation and the role of monetary policy

The prime objective of monetary policy is to promote sustainable economic expansion, for which price stability over the medium run can be considered a precondition. There are, however, differing views on the definition of price stability. In the United States, Federal Reserve Board Chairman Alan Greenspan uses a well-balanced but somewhat vague definition: price stability is achieved when general price developments do not influence the decisions of economic agents.[20] In Europe price stability has been defined more precisely as an inflation rate between 0 and 2%, measured against the relevant consumer price index.[21] In line with this definition, the Nederlandsche Bank has specified price stability as "an inflation not in excess of 2%".[22] In other continents, too, the focus is on consumer prices. New Zealand has announced that "12-monthly increases in the CPI of between 0 and 3% will be considered consistent with price stability".[23] Similarly, Canada has targeted its monetary policies at a rise in the consumer price index of between 1 and 3% and Australia aims for an average underlying inflation rate of between 2 and 3% over the business cycle.[24]

From a purely theoretical perspective, this monetary policy focus on consumer prices has the drawback that asset prices seem to be made irrelevant. A quarter of a century ago, Alchian and Klein (1973) pointed at this shortcoming by arguing that a correct measure of inflation should also take asset price developments into account, to the extent that these determine future consumer prices.[25] More recently, Goodhart (1995) has echoed this argument and has called upon monetary authorities to give asset prices an explicit role in the policy making process in order to prevent the impact of monetary policy on asset markets from accentuating the business cycle.

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20 Greenspan was not the first to use such a definition: his predecessor, Paul Volcker, described "reasonable price stability" in the early 1980s as "a situation in which ordinary people do not feel they have to take expectations of price increases into account in making their investment plans or running their lives" (Volcker and Gyothen (1992), p. 178).

21 In a report to the EU central bank governors, the Group of Experts chaired by R. Raymond (1990) defined price stability as "a level of inflation close to zero [...] i.e. a maximum of 2% in the medium run".

22 De Nederlandsche Bank, Annual Report 1994, p. 21.

23 As specified in the January 1997 Policy Targets Agreement between the minister of finance and the central bank governor.

24 The inflation objective in Australia relates to the headline consumer price index, excluding special factors that are considered either very volatile (such as fresh fruit and vegetables) or not directly related to domestic demand pressures (such as changes in the mortgage rates or in the prices of certain public services).

25 More specifically, Alchian and Klein (1973) argue for the construction of a "constant utility" price index using *inter alia* futures prices for all relevant goods and services. When such prices are not available, asset prices can act as substitutes as these reflect the current price of future consumption flows.

However, this viewpoint creates more difficulties than it resolves. If the end objective for monetary policy were to be broadened beyond consumer prices by focusing on some amalgamated index that also included asset prices, this would, in practice, create new problems of its own. Given a much higher volatility of asset prices than consumer prices, targeting the stability of this index could be expected to lead to greater and more frequent adjustments in monetary policy, which would have adverse consequences for the stability of consumer prices and output.[26] Viewed from this perspective, it is questionable whether asset prices should play a substantive role in the determination of monetary policy.

The crux of the problem lies in the fact that developments in asset prices may be driven by changes in many more or less "fundamental" factors – such as expected rates of return, time preferences, fiscal treatment, or risk premia – that in principle need not prompt an adjustment to the monetary policy stance. This makes clear that the main difficulty for policy makers is the identification of asset price inflation. The foremost difficulty lies in establishing whether an asset price development should be attributed to real or inflationary pressures. But, as illustrated by equation (1), there is a further complication to the extent that policy may need to react differently to a price change towards, rather than away from, equilibrium. If an asset price increases, but the price level is brought closer to equilibrium, a policy reaction would be destabilising. The identification problem is thus twofold: first, in finding out to what degree an asset price change reflects real factors and, second, in identifying how the new price relates to the equilibrium price on the relevant asset market.

Moreover, at a technical level, the construction of a relevant asset price index is problematic. As the asset market consists of numerous sub-markets with generally heterogeneous products, changes in expenditure patterns are relatively pronounced and differences in product quality have a relatively strong impact on price developments. As a consequence, it is hardly possible to construct a representative asset price index. Although these measurement problems also apply to the consumer price index, they are significantly smaller: consumer products are more homogeneous, the pertinent expenditure patterns are less variable, and the time horizon determining the value of these products is much shorter.

4.3 The impact of asset price inflation on consumer prices and output

A more pragmatic approach to the question whether asset price developments should influence monetary policy is by determining how these may affect the (current and future) stability of consumer price inflation or output growth. In this context, along the lines of the analysis set out in the previous sections, four main channels can be identified through which asset price developments influence inflation or output.

First, asset price developments may have a direct impact on consumer prices to the extent that they prompt adjustments in the prices of services of capital goods (for instance, rents will rise when the underlying value of real estate increases). In view of the lags and rigidities constraining such price adjustments, this effect is weak in the Netherlands. Second, changes in the value of assets may lead to adjustments in domestic expenditure, which in turn may indirectly affect output growth and inflation. As set out in Section 3.3, these wealth effects do not seem very large in the Netherlands; this is confirmed by a relatively low estimated elasticity in the MORKMON-model of the Nederlandsche Bank.[27] Third, asset price changes may spawn confidence effects, which will also indirectly impact on the level of domestic expenditure (on both consumption and investment products). Fourth, asset

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[26] The extent to which this would occur depends on the weight assigned to asset prices in the combined index.

[27] The elasticity of private consumption to household wealth is estimated at 0.05 in this model, described in Fase, Kramer and Boeschoten (1992); see also footnote 16. Strictly speaking, the third and fourth channel set out hereafter also influence domestic expenditures; however, it is unclear to what extent this elasticity fully reflects these channels, particularly in exceptional circumstances (such as those prevailing in Japan in recent years).

price developments may generate a self-reinforcing effect through the credit channel, thereby creating a procyclical influence on the business cycle. More specifically, a rising asset market will boost the value of available collateral, which in turn may fuel the growth of credit and spending (including that in the asset market). Conversely, when an asset market loses value and individuals experience difficulty in servicing their debts, banks will suffer losses and may react by reducing credit, thereby magnifying the cyclical downturn (through a "credit crunch").

This analysis of the relationship between asset prices on the one hand and consumer prices on the other, is instrumental in determining whether monetary policy should react to asset price developments. To the extent that consumer prices are directly influenced by asset price increases or are indirectly subjected to upward pressure as a result of adjustments in domestic expenditures (reflecting wealth and confidence effects), some monetary policy tightening will – other conditions being equal – be appropriate to maintain price stability. Conversely, asset price declines will reduce upward pressure on consumer prices, thereby creating scope for a monetary easing to support output growth. This indicates that pronounced asset price developments may, in principle, influence the monetary policy stance, even when this policy is primarily focused on (developments of) consumer prices.

However, monetary policy may not be the preferred instrument in reaction to the impact of asset price inflation through the fourth channel – the provision of credit. In particular, if the previous three channels are weak and asset price inflation takes place in a context of relatively stable (current and expected) consumer prices, the risks will mainly relate to the prospective stability of the financial institutions. In this situation, asset price developments are likely to represent a greater risk for the future stability of output growth than that of (future) inflation. The recent Japanese experience is a case in point.

In effect, when a self-reinforcing interaction takes place between the development of bank credit and asset prices, but consumer prices remain stable, monetary policy is a rather blunt instrument as it influences many other variables besides this interaction. (Assuming, of course, that this instrument is implemented in a market-friendly manner within a liberalised financial context). A more targeted instrument to address a credit spiral is supervision policy, which can contribute to a sufficiently prudent implementation of credit standards by banks, for instance through collateral requirements that offer adequate protection against asset price collapses. To the extent that banks use the asset price and business cycle-upswings to make provisions for less favourable times – as responsible banks should – the credit channel will not have a procyclical influence. Using this instrument kills two birds with one stone: more emphasis on prudential policies will enhance prospects not only for continued financial stability, but also (by dampening banks' credit activities) for sustained price stability. More generally, as set out in the so-called Tinbergen rule, aiming for two objectives (price stability and financial stability) requires at least two instruments (interest rate policy and prudential policy).

In practice, recent developments in Dutch mortgage lending also call for prudential vigilance, since banks have eased the conditions under which they provide mortgage credits (in particular as noted in Section 3.2, both partner and temporary incomes can now be fully counted in determining the credit ceiling; many banks have also increased the level of this ceiling relative to the value of the underlying real estate). A further warning is provided by banks' greater marketing efforts for these credits, for instance by means of fiscally attractive packages.

4.4 Asset prices as leading indicators

When monetary policy is primarily aimed at consumer prices, the role of asset prices will, in practice, be subsidiary and may thus be small. However, asset price developments could also play a part in the monetary policy process to the extent that they provide leading information on future consumer price and output developments. Indeed, studies confirm that asset prices do have some leading indicator properties. However, for the Netherlands the results of these studies, particularly regarding prospective inflation, are not unambiguous and are subject to major uncertainty ranges that vary over the course of time. Specifically, using bivariate regressions, Borio et al. (1994) establish a positive relationship between asset and consumer price developments, but do not find this relationship to be statistically significant. Bikker and Kennedy (1997), by contrast, establish a negative relationship between the deviations from trend of asset and consumer price developments; however, this relationship is relatively weak and is the result of two simultaneous influences between these two variables (the dominant influence is that rising asset prices generally reflect a lowering of inflation expectations which – when the expectations turn out correct – later translate into a moderation of consumer prices). In all, asset prices do not seem to possess particularly reliable information, thereby limiting their usefulness for monetary policy purposes.

Conclusions

In summary, there are valid reasons to focus monetary policy primarily on consumer price developments; this enhances monetary instrument stability and precludes the considerable identification problems of asset price inflation. In this situation, asset price developments play a role to the extent that they directly influence consumer prices or indirectly lead to changes in domestic spending through wealth and confidence effects. In addition, asset prices may influence monetary policy on account of their leading information for future consumer price and output developments; however, in practice this information does not appear to be very reliable and may thus best be used as a complement to other leading indicators. When asset price inflation interacts with credit growth, supervision policy can help contain risks of future financial instability while also assisting monetary policy efforts aimed at maintaining consumer price stability.

More specifically for the case of the Netherlands, where the monetary strategy has been consistently anchored in an exchange rate commitment over a period of many years, the scope for making active use of monetary instruments for domestic (asset price or other) objectives is extremely limited. This constraint at the macro-level adds to the arguments to address exuberant asset price developments with policy instruments at the micro-level. This does not, however, resolve the problem that it is well-nigh impossible to conclusively identify excessive asset price changes in practice. Supervision policies will thus need to be especially on the alert in a buoyant market in order to prevent asset prices from having a destabilising influence, through the credit channel, on medium-term prospects for sustained non-inflationary growth.

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Equity prices and monetary policy in the United States

Vincent Raymond Reinhart*

Introduction

The wide swings in equity prices on world markets in recent months have riveted investors' attention. With more than $11 trillion of equity wealth in the hands of the US public, such investor interest is understandable. A more difficult question is whether equity prices should receive as much prominence in the deliberations of central bankers as they do in the financial press. To the extent that equity prices are sensitive to real interest rates and contain the market's implicit assessment of corporate profits going forward, they could be read as signaling the current and future state of the economy. They could also directly affect spending by their contribution to wealth and influence on relative returns. However, that equity price *could* be important in either signaling or directly influencing the ultimate objects of a central bank's intentions – economic activity and inflation – does not necessarily imply that monetary policy *should* be sensitive to them.

To examine the case for responding systematically to stock prices, I consider three questions:

- First, how are equity prices related, even if quite imprecisely, to macroeconomic fundamentals?
- Second, do equity prices systematically influence macroeconomic outcomes?
- Third, does attention paid to the equity market on the part of monetary policymakers feed back on the dynamics of economic activity, inflation and equity prices?

The first question is related to the determinants of equity prices, a subject examined in Section 1. I consider a variety of simple valuation equations that link stock prices to interest rates and economic conditions. All those relationships convey the same general impression: equity prices are volatile relative to fundamentals and are currently on the high side of that predicted by history.

As to the second question, based on another set of estimations, this time using quarterly data on real GDP and some of its components, I show in Section 2 that equity prices importantly influence spending. Moreover, that wealth effect is more evident in the past twelve years than it was in the prior twelve years. Proportionally, much of that effect on GDP is recorded in its investment component, although consumption is estimated to be sensitive to stock market wealth as well. Offsetting the effect on aggregate demand somewhat, real imports are also sensitive to equity prices.

The two sections that follow address the third question. From a monetary policy perspective, a central bank even with a purely macroeconomic objective – say, containing inflation pressures – must be sensitive to equity prices. But, in that regard, a forward-looking central banker has a responsibility to monitor any financial variable thought to influence economic activity and pressures on inflation, including the foreign exchange value of the dollar and nominal and real interest rates all along the term structure. Elevating equity prices beyond the function they serve in helping to predict the ultimate objects of central bank concern could pose problems. For example, a simple theoretical model of the economy suggests that responding to the *level* of equity prices raises the net effect of news on equity prices (Section 3). Should the central bank respond to the *change* in equity prices, it may well raise the volatility of share prices in response to monetary policy misalignments and, depending on parameter values, actually destabilize the economy (Section 4). Concluding comments are provided in the final section.

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1. Equity price fundamentals

All modern asset pricing relationships begin from the marginal condition that households substitute consumption over time and allocate their assets at a point in time so that risk-adjusted expected rates of return are equated to the risk-free real interest rate.[1] If P is the price of goods and services, Q the price of an equity share D, the dividends provided per period, and r the risk-free real rate, then:

$$r_t = \frac{D_t}{Q_t} + E\frac{Q_{t+1}/P_{t+1} - Q_t/P_t}{Q_t/P_t} - \rho_t$$

where E denotes the expectation conditioned on all available information. In this relationship, ρ is the equity premium, which depends on how returns covary with the marginal utility of consumption and is not presumed to be constant. All rates of return refer to the period over which households can change their consumption. Similarly, dividends are paid out over that interval. In continuous time models, consumption can vary instant by instant, so that the appropriate real rate is the instantaneous one – effectively an overnight rate.

An analyst could make specific assumptions about the movement through time of dividends and the real short rate to derive an association between the level of D/Q and a market rate. With the help of an auxiliary relationship describing dividend payout policy, or how dividends are related to earnings, Π, a description of the earning-price ratio could be developed. But any such relationship is conditional on those assumptions and holds other variables constant in the background. They are not structural in the way that the intertemporal substitution condition is, implying that they can vary over time. Moreover, different assumptions could even change the market rate that is the appropriate benchmark.

To see this property, I can use the arbitrage condition to solve for the level of equity prices in the absence of uncertainty.[2] The simplest of all pricing models assumes that the real rate, inflation, and the equity premium are all constant and that dividends grow at a constant rate in nominal terms, G. This nominal dividend growth can be divided into its real and inflation-compensation component, as in:

$$G = g + \pi$$

As a result of these assumptions, the pricing formula simplifies to:

$$Q_t = \sum_{i=0}^{\infty} \frac{D_t(1+G)^i}{(1+r+\rho+\pi)^{i+1}}$$

where the exponent in the numerator and denominator differ by one because dividends are assumed to be paid at the end of the period. As long as dividends grow more slowly than the net rate of discount, this summation is bounded and the D/Q ratio can be written succinctly as:

$$D/Q = r + \rho + \pi - G$$

which is often referred to as the Gordon equation, after Myron Gordon (1962).

This simple equation might be served in a variety of flavors. For instance, because dividends grow with the prices of goods and services, this could as well be written in terms of real rates,

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1 These are discussed, for example, in Shiller (1981) and Campbell and Shiller (1987). An extensive review of this work and a discussion of its relevance to the current situation is provided by Cochrane (1997).

2 Exactly why there would be an equity premium in a certain world is a puzzle in itself, but assume that ρ remains. This eliminates all manner of complications associated with nonlinearities and covariance terms but offers another reason why the final pricing equation might be inadequate.

$$D/Q = r + \rho - g$$

Or, define the nominal short rate as R, which sums the real rate and expected inflation. As a result,

$$D/Q = R + \rho - G$$

Or, split the equity premium into a part related to consumption covariation, ρ^c, and another part related to the bankruptcy risk of firms, ρ^b. The sum of the risk free nominal and the bankruptcy-risk term might be proxied by a private nominal rate, R^b, as in:

$$D/Q = R^b + \rho^c - G$$

Lastly, assume that dividends are paid out as a constant fraction, k, of earnings, Π. This implies that dividends and earnings grow at the same rate. This payout rate can transform the constant-growth model to:

$$\Pi/Q = \frac{r + \rho - g}{k}$$

Thus, with not much work, it is possible to derive several variants of a "fair pricing" model. In principle, they all should produce the same predictions. But in practice, it is not likely that analysts will be internally consistent about their assumptions for the right-hand-side terms. More complicated models allow movement through time of dividends, the real short rate, inflation, and the payout rate. But because they are all arbitrary to some degree, they need not produce even roughly comparable estimates of the "fair value" of equity prices. One way to distinguish these alternative explanations of equity prices would be empirical fitness. To that end, I estimated various versions of the Gordon equation using monthly data from 1980.

In principle, the yield on equities should depend on the rate on a competing asset and variables that help to forecast earnings growth and future dividend payouts.[3] The latter category would include inflation expectations (presumably at a long horizon) and the unemployment rate.[4] As for the rate on competing assets, I considered six alternatives, varying by maturity and riskiness. They are the: (1) federal funds rate, (2) three-month Treasury bill rate, (3) ten-year Treasury note yield, (4) three-month commercial paper rate, (5) yield on Moody's AAA-rated seasoned corporate issues, and (6) yield on recently offered A-rated utility bonds.

The Gordon equation relates the level of the return on equities to an interest rate and perhaps other variables. However, standard test statistics of regression equations are only appropriate for stationary variables. Thus, some work may be required to render the underlying time series stationary, if as some researchers, contend, nominal, and even real, interest rates have a unit root (Rose (1988)). In a univariate case, a series with a single unit root may be rendered stationary by first differencing to make it an appropriate regressor or regressand. In a multivariate case, some linear combination of nonstationary variables may be stationary, in which case those variables are said to be cointegrated, in the manner described by Engle and Granger (1987). The appropriate way to combine variables can be imposed from

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3 In recent years, US firms have clearly shifted away from paying dividends, with the ratio of dividends to earnings falling to under 40%. While the theoretical model emphasized what investors receive, dividends, I will concentrate on the earnings-price ratio because of this evident change in firm behavior. But it should be noted that any puzzle of equity price overvaluation would be more intense when the model is written in terms of dividends rather than earnings.

4 Survey measures of longer-term inflation expectations are only available sporadically in the early 1980s. I interpolated values for the University of Michigan's survey of households' five-to-ten-year-ahead inflation expectations using the predictions of a regression of the long-term expectation against monthly readings on one-year-ahead inflation expectations and lagged actual inflation (which are available without gaps).

theoretical priors, estimated by a first stage regression, or calculated through maximum likelihood techniques.

To uncover the specific time-series properties of the variables of interest, I first conducted a battery of tests, which are reported in an appendix, for the presence of one or two unit roots in monthly data for a collection of interest rates and macro variables. It turns out that the levels of nominal interest rates, inflation expectations, and the unemployment rate are all nonstationary. Further, measures of the real rate, or the nominal rate less the appropriate maturity inflation expectations, are also nonstationary.[5] As a result, it is appropriate to allow for the possibility that a coefficient on inflation expectations is not unity and that the relationship explaining nominal interest rates includes other variables. Moreover, it is important to estimate with a technique that generates robust standard errors.

An error-correction model is one simple technique that exploits the dynamic movements of the explanatory variables to estimate a single cointegrating relationship.[6] In estimating an equation that imposes a long-run relationship on the data, I can introduce lagged and led values of changes in the dependent variable and lagged values of the difference between the level of the variable and that predicted by fundamentals, or the error-correction term. Both lags and leads are included because investors are presumably forward looking in their determination of asset values. What is not evident in the technique is how many leads and lags to include. In principle, various combinations of lagged error-correction terms and the led and lagged dependent variable can yield an equivalent representation.

Table 1

Estimates of equations predicting the earnings-price ratio for the S&P 500

Parameter estimates and summary statistics (using monthly data over the period 1980–96)

	Federal funds rate	**Three-month Treasury bill rate**	**Ten-year Treasury yield**	**Three-month CP rate**	**Moody's AAA-rated yield**	**A-rated utility yield**
Speed of adjustment (λ)	-0.082 (0.024)	-0.086 (0.024)	-0.098 (0.025)	-0.084 (0.025)	-0.099 (0.026)	-0.106 (0.026)
Constant	1.962 (1.919)	1.521 (1.791)	0.253 (1.494)	1.326 (1.828)	-0.742 (1.450)	0.256 (1.377)
Interest rate *j*	0.642 (0.133)	0.768 (0.150)	0.973 (0.163)	0.690 (0.140)	1.062 (0.177)	0.940 (0.145)
Inflation expectations	0.167 (0.347)	0.207 (0.322)	0.469 (0.248)	0.198 (0.333)	0.496 (0.242)	0.336 (0.243)
Unemployment Rate	-0.116 (0.264)	-0.137 (0.252)	-0.610 (0.259)	-0.089 (0.256)	-0.784 (0.274)	-0.702 (0.248)
Lagged dependent variable	0.129 (0.069)	0.117 (0.069)	0.088 (0.071)	0.124 (0.070)	0.096 (0.071)	0.080 (0.071)
Led dependent Variable	0.125 (0.069)	0.117 (0.070)	0.096 (0.069)	0.124 (0.070)	0.102 (0.070)	0.089 (0.070)
R^2	0.159	0.164	0.180	0.159	0.175	0.184
Standard error	0.336	0.335	0.332	0.336	0.333	0.331

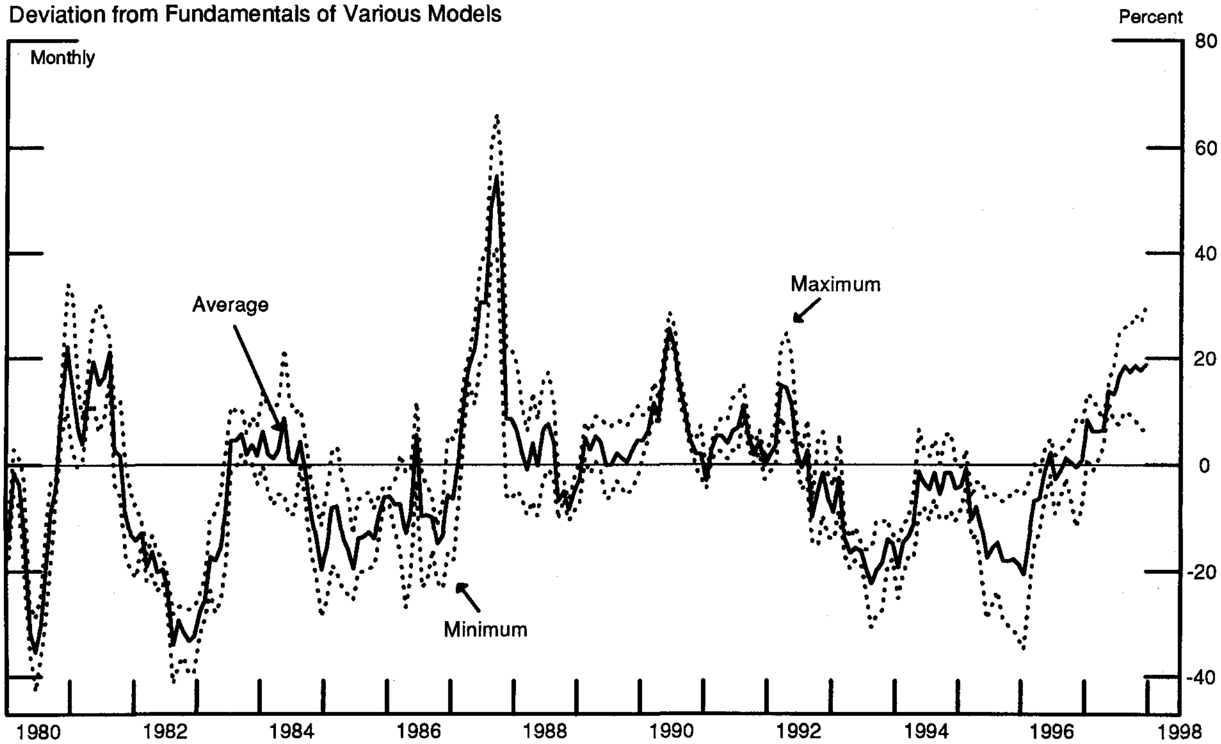
Note: Standard errors in parentheses.

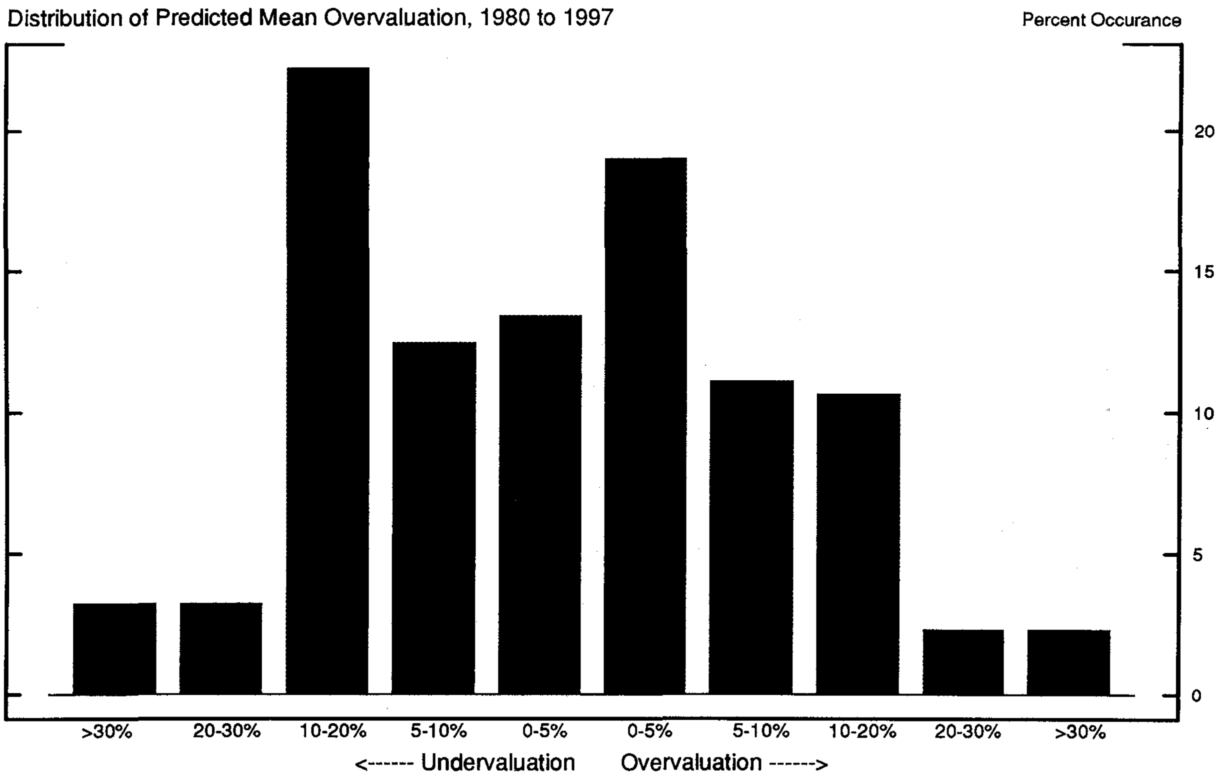
5 At a broader level, it may be difficult to imagine an economy where both nominal and real interest rates are nonstationary, particularly, as Cochrane (1991) points out, when those variables move in a narrow range over very long samples.

6 A discussion of various techniques is provided in Campbell and Perron (1991).

Figure 1

Predicted overvaluation of S&P 500





To keep matters simple and to avoid estimating many parameters, I examined a specification of the form:

$$\Delta\frac{\Pi_t}{Q_t}=\frac{\lambda^j}{k}\sum_{i=1}^{k}\left(\frac{\Pi_{t-i}}{Q_{t-i}}-a_1^j-a_2^j R_{t-i}^j-a_3^j\,\pi_{t-i}^{10}-a_4^j u_{t-1}\right)+b^j\,\Delta\frac{\Pi_{t-1}}{Q_{t-1}}+c^j\Delta\frac{\Pi_{t+1}}{Q_{t+1}}+\eta_t^j$$

one for each of the alternative right-hand-side variables. Taking the simple average of the lagged error-correction term follows the advice of Berkowitz and Giorgianni (1996) and permits an easy exploration of alternative lag lengths (the parameter k) while keeping the number of parameters estimated constant. Notice that this is an appropriate object of estimation because the terms in the summation should be stationary if a cointegrating vector exists and the other variables, both on the left and right-hand side of the equation, are stationary because they appear in difference form. If the speed of error correction, the coefficient λ, is significantly negative and either or both a_2 and a_3 are nonzero, then a cointegrating vector exists.

After some experimentation with lag length, these six equations were estimated individually with two lags of the error-correction term using monthly data from 1980 to 1996. Those regression results, summarized in Table 1, indicate that the earnings-price ratio moves positively, and mostly one-for-one, with nominal interest rates. Among those interest rates, private longer-term yields seem to have slightly more predictive power. There is an important business-cycle element in pricing, in that the coefficient on the unemployment rate is usually significantly negative. However, somewhat inexplicably, expected inflation plays no role in these estimates. In all cases, the speed of error-correction is estimated to be significantly negative, implying that earning-price yield tends to revert to a moving mean determined by fundamentals. That is, because the level of the nominal rates on the right-hand side helps to explain the change in the earnings-price ratio over time, the earnings-price ratio is cointegrated with these nominal rates (individually).

I evaluated those fundamentals for the six alternative specifications from 1980 onward, using only the level portion of the estimation equation and not the dynamic terms. The percent deviation of the predicted earnings yield based on macro variables each month from the actual earnings yield gives an estimate of the under- or overvaluation of the S&P 500, the range for which is provided in the upper panel of Figure 1. The dotted lines give the maximum and minimum predicted deviation from fundamental valuations, judged each month across the six models, while the solid line gives the mean of those six estimates. As is evident, equity values are considered by these models to be currently overvalued – to the tune of 6 to 28%. But even that excess is dwarfed by the significant bubble in prices, at least as viewed by these models, in summer 1987.

The bottom panel sorts the mean deviation from fundamentals by size. The resulting distribution is distinctly bimodal, producing spikes at significant undervaluations (10 to 20%) and at correct valuations (0 to 5%). Adding to the widespread uneasiness about equity prices reported in the financial press is the swing in fundamentals over the past two years: The models believed equities to be about 20% undervalued at the end of 1995, but the subsequent appreciation of share prices more than eroded that mispricing to produce the current estimated overvaluations.

2. Evidence on the effect of equity values on spending

Asset valuation is only directly relevant to monetary policymakers if it can be determined that those asset values influence aggregate demand. This presents another opportunity to estimate an error-correction model, this time to establish the long-run relationship among financial market quotes and macroeconomic outcomes. Over the years, economists have purported that two sets of market quotes have some predictive power for aggregate demand: the real value of equity prices (as in Sprinkel (1964)) and the slope of the term structure (as in Estrella and Hardouvelis (1991)). Because of this predictive power, both of these variables are in the index of leading indicators. My strategy is to offer both as potential explanatory variables for real GDP and its main components, consumption and business fixed investment, and let the data determine relative merits.

Some analysts have given a structural interpretation to the inclusion of the yield curve in an explanation of aggregate demand. According to Benjamin Friedman (1978), for instance, the long-term nominal interest rate importantly captures inflation expectations, so that the slope of the term structure proxies for the (negative of) the real short-term interest rate. I will not offer that structural interpretation. Rather, I seek to determine the incremental influence of relative equity prices given the presence of an indicator that captures the regularities of the business cycle. As a result, though, the estimates will not be directly comparable to those of other researchers examining the role of wealth in influencing consumption or saving using structural relationships.[7]

Because dynamics are likely to be difficult to disentangle, I estimate a model of the form:

$$\Delta y_t^j = \frac{\lambda^j}{k} \sum_{i=1}^{k} \left[y_{t-i}^j - a_1^j - a_2^j (R_{t-i}^{10} - R_{t-i}^{3m}) - a_3^j (q_{t-i} - p_{t-i}^j) \right] + b^j \Delta y_{t-1}^j + c^j \Delta y_{t+1}^j$$

where y^j represents the logarithms of, respectively: (1) real GDP, (2) real consumption, (3) real business fixed investment, and (4) real imports of goods and services.

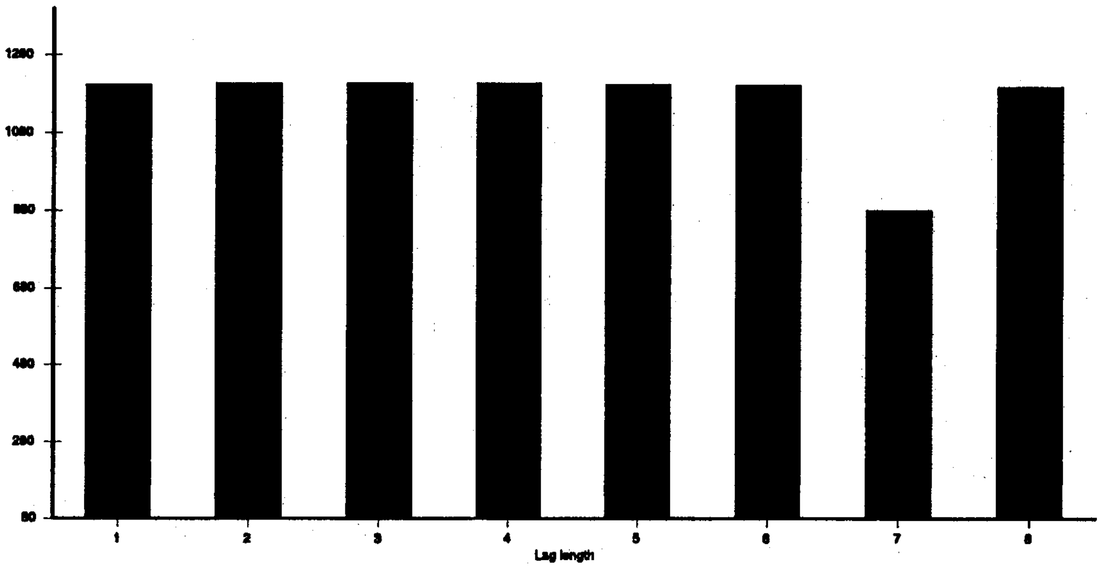
Again, dynamic terms are included to improve the precision of the estimates of the long-run relationship, not to generate an equation with which to forecast near-term behavior. While the yield curve slope is identical across equations, the S&P 500 (q) is deflated by the price index specific to that component of GDP (p^j).[8] As the components of GDP are likely to be highly correlated with each other and the total, I will estimate the four relationships simultaneously over quarterly data from 1973 to 1996 for various lag lengths of the error-correction summation. I chose this subset of the data for which income statistics are available on the hypothesis that observations drawn from the fixed-exchange-rate sample are not relevant for current experience.

Figure 2

Estimates of the model explaining real GDP and its components

Using quarterly data over the period 1973–96

Value of the Likelihood Function



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[7] Deaton (1992) provides a comprehensive review of the consumption literature. A description of consumption in a large structural model applying conventional theory, with an emphasis on shifts in behavior over time, is given in Mauskopf (1990).

[8] I found, but do not report, similar results to what follows using a broader measure of wealth from the flow-of-funds accounts.

As is evident in the upper panel of Figure 2, a lag of four quarters maximizes the log likelihood function, perhaps suggesting a residual seasonality in the data.[9] On the possibility that the structure of the model is not homogeneous through time, I re-estimated over two samples, from 1973 to 1984 and from 1985 to 1996, which represents an equal division of the data. As is evident from Table 2, the data are clearly drawn from (at least) two distributions. Pre-1985, financial market variables do not appear to exert a systematic effect on spending. Post-1985, both relative returns and the value of equity prices influence GDP and its components.

Table 2

Parameter estimates and summary statistics

	1973–84				1985–96			
	GDP	C	I	M	GDP	C	I	M
Speed of adjustment (λ)	-0.038 (0.020)	-0.045 (0.012)	-0.027 (0.029)	-0.044 (0.041)	-0.083 (0.024)	-0.095 (0.024)	-0.069 (0.025)	-0.067 (0.040)
Constant	8.743 (0.189)	8.503 (0.122)	6.416 (0.460)	5.848 (0.269)	8.328 (0.055)	7.938 (0.053)	5.611 (0.408)	5.324 (0.303)
$R^{10} - R^{3m}$	0.190 (0.103)	0.158 (0.042)	0.591 (0.669)	0.420 (0.393)	0.027 (0.013)	0.039 (0.013)	0.164 (0.081)	0.109 (0.083)
Relative equity prices	-0.479 (0.238)	-0.612 (0.134)	-0.776 (0.831)	-0.634 (0.563)	0.325 (0.035)	0.321 (0.033)	0.714 (0.237)	1.030 (0.281)
Lagged dependent variable	-0.094 (0.098)	-0.439 (0.098)	0.078 (0.093)	0.099 (0.131)	-0.205 (0.109)	-0.422 (0.098)	-0.105 (0.107)	-0.315 (0.112)
Led dependent Variable	-0.262 (0.104)	-0.182 (0.098)	-0.025 (0.101)	-0.126 (0.136)	-0.205 (0.109)	-0.389 (0.099)	-0.303 (0.105)	-0.302 (0.115)
R^2	0.301	0.499	0.484	0.257	0.090	0.193	0.205	0.127
Standard error	0.010	0.006	0.023	0.037	0.005	0.005	0.016	0.018

Note: Standard errors in parentheses.

For whatever reason, over the past twelve years, the real value of equities (appropriately deflated) appears to exert a significant and sizable positive effect on spending. The elasticity of aggregate spending to stock market prices, at about $\frac{1}{3}$, is a weighted average of an elasticity of about $\frac{1}{3}$ for consumption, over $\frac{2}{3}$ for investment, and 1 for imports, where the weights are their contributions to GDP. Thus, a rise in equity prices spurs spending on goods from abroad more than at home, limiting the net impact of that wealth gain on real GDP. As to domestic spending, because the effect of equity prices appears to be greatest for investment, rising equity prices have no doubt importantly contributed to the expansion of capital in recent years. Similarly, investment has a more cyclical response to the slope of the term structure, with a semi-elasticity of about $\frac{1}{6}$ for each percentage point change in the spread of the ten-year yield over the three-month bill rate. In all four cases, the speed of error correction is significantly negative, indicating that I have identified cointegrating vectors.

The net result of this estimation exercise is to suggest that spending, both total GDP and its major components, are related to equity prices, but that these relationship are sensitive to the period chosen. Because of their influence on aggregate demand, central bankers have some grounds for consulting equity prices in the formulation of policy.

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9 As is evident in the figure, the likelihood function for this estimation (and the others that follow) tends to be quite flat. Unfortunately, the parameter estimates – and policy implications – differ across the lag lengths chosen, even though they have fairly similar explanatory power. One reason researchers may differ in their enthusiasm for various indicators may be that they are sampling in different regions of this flat likelihood surface.

3. Monetary policy and equity prices

Monetary policymakers might be concerned about equity prices straying above fundamentals because of potential adverse effects as prices rise or as they fall back to earth.[10] As to the rise, if prices are moving above fundamentals, relative prices are misaligned, dictating some misallocation of resources. Households might be consuming out of their paper wealth, firms buying capital based on inflated market relative to book values, households and businesses taking on debt because leverage ratios look good, and new firms starting up because capital markets are so receptive.

Monetary policymakers might also be concerned about increases in stock prices because of a fear that there was time dependence to overvaluation – on the logic that the longer prices stray above fundamentals, the further they stray and the harder they will fall. Equity prices might matter on the way down because of concerns about systemic risk, knock-on effects on spending and confidence, and the risk of subsequent undershooting. However, mechanisms are well developed for dealing with systemic risk, including lending via the discount window and a willingness to add ample reserves at times of stress.

A simple model

A simple two-equation dynamic model can capture the linkage between monetary policy and equity prices and provide an example of a near-rational bubble. It also suggests caution in thinking that policy action can automatically deflate an asset bubble. The model is a variant of that first provided by Blanchard (1981) but which has since been employed by Blanchard and Dornbusch (1984), Branson, Fraga and Johnson (1985), and Dornbusch (1986), among others, to examine a variety of policy issues.

To be specific, suppose that aggregate demand, y, depends on the real short-term interest rate, r, and relative equity prices, q, as in:

$$y = a_1 - a_2 r + a_3 q$$

I define the real rate as the difference between the nominal rate, i, which is set by the central bank, and instantaneous inflation, π.[11] Equity prices may enter in determining aggregate demand because of wealth effects on consumption or because Tobin's q influences investment spending.

As to the rest of the economy, inflation follows inertially from previous excesses of spending over potential output, k, so that:

$$\dot{\pi} = b(y - k)$$

where the dot over a variable signifies the time derivative. This is an accelerationist-style Phillips curve implying that the current level of inflation is a backward-looking average of past inflation.[12]

Monetary policy smooths the nominal interest rate according to the output gap and inflation, following the insight of John Taylor (1993) that Federal Reserve behavior of the past decade could be related to the recent behavior of measured inflation and an estimate of the excess of output over its potential. While Taylor specified a relationship explaining the level of the federal funds rate, the proximate instrument of monetary policy in the United States, an equation describing the change in that rate would also fit the data, as well as the stylized fact that policy rates tend to evolve smoothly through time. The exact relationship is given by:

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[10] A discussion of the role of a central bank confronted with an asset bubble is provided by Miller (1996).

[11] Lower-case variables represent logarithms and I will suppress time subscripts wherever possible.

[12] A forward-looking inflation process of the sort modelled by Calvo (1983) would alter the results because it introduces fundamental problems of indeterminacy, some of which are discussed in Reinhart (1992).

$$\dot{i} = \alpha(y - k) + \beta\dot{\pi}$$

Further assume that, in this first-difference form of the Taylor rule, the net responsiveness to the change in the inflation rate is unity ($\beta = 1$), with the result that the change in the real rate is a function of the output gap. The consequence of such a policy rule combined with the backward-looking specification of inflation determination is that the central bank is merely satisfied with making the inflation rate follow a random walk, and the real rate follows:

$$\dot{r} = \alpha(y - k)$$

Some of the issues involved with the central bank following a rule conditioned only on real variables, that is, lacking a nominal anchor, are examined in Reinhart (1991). But for the purposes at hand, this rule serves to establish the important linkages between equity prices and monetary policy.

Equity prices satisfy the intertemporal arbitrage condition discussed in the previous section, written in continuous time as:

$$r + \rho = \frac{\delta}{q} + \frac{\dot{q}}{q}$$

where δ denotes the fixed stream of dividends and ρ is the premium required of equity investments; both the dividend rate and the equity premium are assumed to be constant, although it is of no consequence, except in complicating the notation, to make them functions of income.

Some steady-state comparisons

The solution to this model is straightforward. In the long run, output must equal its potential for policy to be at rest,

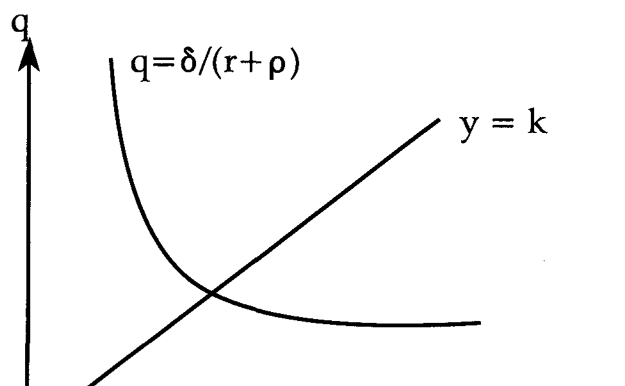
$$k = a_1 - a_2 r + a_3 q$$

and relative asset prices must satisfy the familiar rectangular hyperbola,

$$q = \frac{\delta}{r + \rho}$$

The determination of the real rate and equity prices, as in Figure 3, is set at the intersection of the two schedules.

Figure 3



Even at this basic level, before introducing dynamics, the model offers insight as to two structural changes that could bolster equity prices. As a first possibility, note that the level of potential output fixes the position of the spending-balance relationship. If the economy can produce at a higher level of resource utilization without generating inflationary pressures, real interest rates must be lower in the long run. But a lower rate of discount supports higher equity values, which would be seen as the movement up the equity-price-real-rate hyperbola. Notice, however, that the rise in equity prices should, in principle, be explained by lower real rates. Thus, the general failure of statistical models to predict the runup in share values over the past few years, along with the apparent continued elevated levels of real interest rates, suggests that more is at work than an increase in the economy's potential.

As a second possibility, note that the equity-price determination equation included ρ, the equity premium. That term is necessary to account for the fact that the mean excess return on equities over a risk-free rate in the United States is typically estimated at around 5 to 6% over long periods (as first documented by Mehra and Prescott (1985)). This equity premium is a puzzle in most theoretical models, in that it would be significantly eroded if households exhibited just a little willingness to shift their consumption intertemporally. Instead, the relatively high historical return on equities given the smoothness of consumption would seem to imply that consumers inordinately value stable consumption. In the simple model, a reduction in the equity premium shifts the long-run equity price locus outward, along the fixed policy rule, and is consistent with both a higher real interest rate and higher equity prices. From this perspective, the recent string of high returns could be seen as the price realignment required to pare the expected equity premium.

Many reasons might be offered as to why the equity premium has declined, including increased concerns about the value of social security benefits, the spread of defined-contribution pension plans, declines in the transactions costs and enhanced publicity of mutual funds, the aging of baby boomers, or even a decline in the expected volatility in nominal and real returns as the Federal Reserve makes further progress toward price stability. Whatever the reason, an exogenous decline in the equity premium would raise both equity prices and the real rate. Moreover, statistical models of the type estimated in the previous section would fail to explain that configuration of market returns because the regime over which those models were estimated no longer held sway. However, it is somewhat risky to emphasize an explanation that required jettisoning a regularity describing behavior of at least the past century. Further, this is less an explanation than a rationalization, because it pushes the fundamental explanation of high equity prices, why ρ declined, outside the scope of the model.

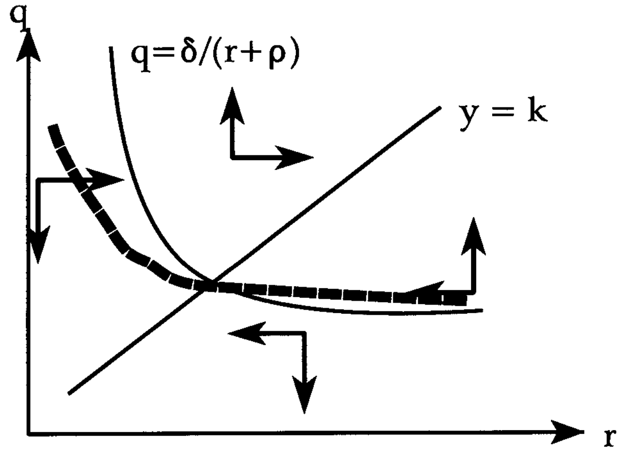
Dynamics

The movement from steady state to steady state can be described by two equations of motion (the rate-smoothing and equity-price-arbitrage relationships) which pinpoint r and q at each moment. As shown in Figure 4, there is a unique set of equity prices and real interest rates that provides for a stable transit to the steady state. Those points slope downward along the dashed line. However, any point in that quadrant can set off a dynamic path that satisfies all the equations of the model. It is only those points along the downward sloped saddlepath that satisfies the model and returns the system to its steady state from a point different from the long-run equilibrium.

Those points along the saddlepath match intuition: if monetary policy were unexpectedly loose, that is, if r fell below its long-run level, equity prices would jump up on impact. The spur to spending associated with a lower real rate and higher equity prices would send output above its potential, leading monetary policy to tighten gradually to return the economy to its potential. The relationship lurking in the background, the Phillips curve, implies that inflation will be permanently higher in an amount depending on the cumulative upward deviation of output from its potential. It is in this sense that high share price might be seen as emblematic of policy laxity. This view has led some, going back at least to Sprinkel (1964), to advocate that the central bank accord asset prices some priority among the indicators that it consults.

Regardless of the broader merit of that argument, four points should be considered before translating generally high equity values into a specific indictment of Federal Reserve policy. First, the

Figure 4



support provided by low interest rates to equity values should be shared by all long-lived assets. As yet, there is no evidence of widespread asset-price inflation, in that land, home, and commodity prices appear well contained. Second, the key mechanism by which policy bolsters the stock market is by pushing the real rate below its equilibrium level. While both the real short rate and its equilibrium are, to varying degrees, unobservable, reasonable proxies of current real rates are not particularly low. Third, within this tiny model, the Phillips curve rules, and temporarily high equity prices should be associated with a quickening of inflation. No such pressures are evident as yet. And fourth, at no point in the dynamic adjustment triggered by inappropriately easy monetary policy could equity prices be said to be overvalued. Rather, it is low real rates that buoy their value, a factor that should be captured in statistical models of equity values.

A near-rational bubble

In this model, an asset bubble would entail increases in equity prices beyond fundamentals that still allowed all relative prices to satisfy investors' arbitrage condition, spending balance, and the policy rule.[13] For instance, suppose that equity prices were just a bit higher than the long-run equilibrium level but the real rate was at its natural level, as at point A in Figure 5. Such a point could be thought of as a situation where monetary policy was correctly aligned for the steady state but equity markets misaligned, or monetary policy was currently too loose for the current level of equity prices. That point would be an equilibrium because expectations of future changes in equity prices would make both investors and policymakers content with all asset prices. The problem is that this starting point sets off explosive dynamics where equity prices are continually rising, as denoted by the thick solid arc. Along that transition, equity prices do not yield excess returns, rather, the low dividend-price ratio is just offset by capital gains.[14]

Monetary policymakers shadow the increase in equity prices by raising short rates, as required by the Taylor rule when aggregate demand rises as the result of higher-wealth-induced increases in aggregate demand. However, the gradualism embodied in the Taylor rule is not sufficient to offset the full extent of the effects of the rise in wealth: Because aggregate demand remains always above potential output, inflation is rising continually during this episode. Thus, the model yields the result that policymakers tighten in response to an asset bubble, but some of the effects of the rise in equity prices spill

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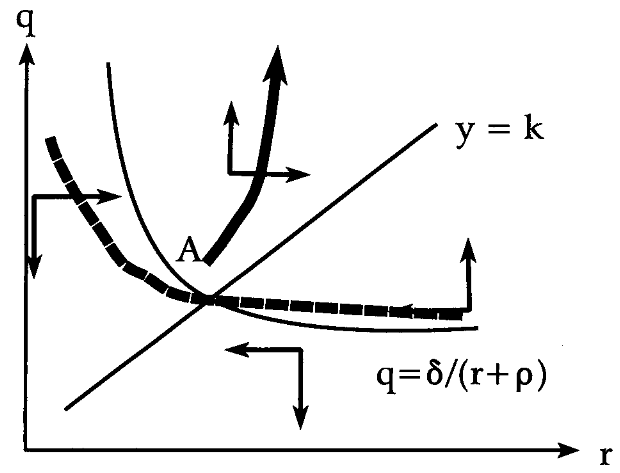
13 Work on asset-price bubbles was pioneered by Flood and Garber (1980). An attempt to determine if equity markets could be characterized as having a rational bubble can be found in Craine (1993).

14 This is a near-rational bubble because investors should rationally expect that at some point it will pop. To make this a rational bubble, the model would have to include, and price, a random hazard that such an event would occur, as in Blanchard and Watson (1982).

Figure 5



over to inflation. Note also that tightening does not pop the bubble. Indeed, along the transition, capital gains are driven by the current short rate. So, *if the central bank raises the short rate, the asset bubble inflates faster to give investors the requisite higher return.*

Were there a rational bubble inflation in the current situation, we would expect that empirical models based on the long-run determination of equity prices would underpredict. This follows because the term dropped in moving from the instantaneous arbitrage condition to the long run, the anticipated change in equity prices, should systematically bias the predictions of empirical models. However, because a bubble does create wealth, we would expect it to generate excess demands for goods and inflationary potential. It is also the case that the evidence for rational bubbles is slim to nonexistent: As yet, no one has found a rational bubble in US history.

4. The systematic response of monetary policy to equity prices

The model also can be used to entertain questions about the appropriate degree to which monetary policymakers should respond to equity prices. Even with the simple Taylor rule, the central bank is responding to equity prices to the extent that those prices are important in influencing spending. That is, substituting the determination of aggregate demand into the policy rule yields:

$$\dot{r} = \alpha(a_1 - a_2 r + a_3 q - k)$$

Notice that in this formulation, equity prices are important for what they imply about future spending and inflation, not for their own sake. Indeed, they have the same status as any of the other critical determinants of spending, production, and inflation, such as the foreign exchange value of the dollar, federal spending and taxes, and foreign economic activity, which are subsumed in the constant terms of this model. But, as already mentioned, equity prices straying from their long-run level directly corresponds to a monetary policy misalignment. This might suggest keying policy choice to share values, if the central bank knew fundamental valuations.

Elevating attention to equities beyond their direct consequences to spending would entail adding another term to this rule. Policymakers might be confident about fundamentals, in which case the central bank could respond incrementally to the *level* of equity prices or, if unsure of the long-run value of the stock market, they might move their policy lever based on *changes* in equity prices.

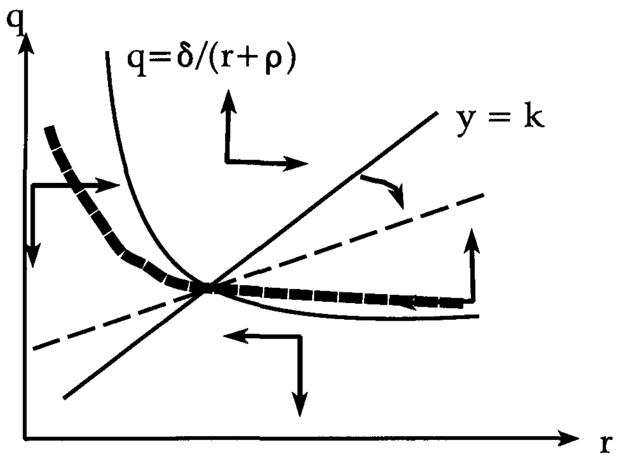
Feedback to the level of equity prices

If policymakers cared about the level of equity prices beyond any consequences for spending, they could modify their rule to:

$$\dot{r} = \alpha(a_1 - a_2 r + a_3 q - k) + \gamma(q - \bar{q})$$

where the line above a variable denotes its steady-state value. This additional concern about equity prices has the effect of flattening the long-run locus at which policy is at rest, as in Figure 6. Essentially, the central bank is no longer satisfied with responding solely to output varying from its potential; it also moves when out of the steady state because share values differ from their long-run levels. This flattens the saddlepath. While changes in equity prices around a steady state are damped, transitions from one to another are more abrupt. Thus, the consequence for volatility is not obvious. If prices move mostly because of the release of information regarding fundamentals, then there would be larger discrete changes. If prices move mostly to generate the capital gains required to equate returns, then volatility would be damped.[15] The profession's general inability to explain equity prices, which was also evident in the regression exercises earlier, would seem to imply that movements in the former, not the latter, camp are importantly responsible for volatility.

Figure 6



Responding to changes in equity prices

Concern about the change in equity prices could be expressed in a policy rule written as:

$$\dot{r} = \alpha(a_1 - a_2 r + a_3 q - k) + \lambda\frac{\dot{q}}{q}$$

This might be justified by the central bank leaning against the wind of surges or collapses in share values. It might particularly find favor because the rule does not presume that the central bank knows the long-run level of equity prices, but rather responds to an observable, for example, recent changes in

15 A similar ambiguity in the effect of capital controls on asset-price volatility is shown in Reinhart (1998).

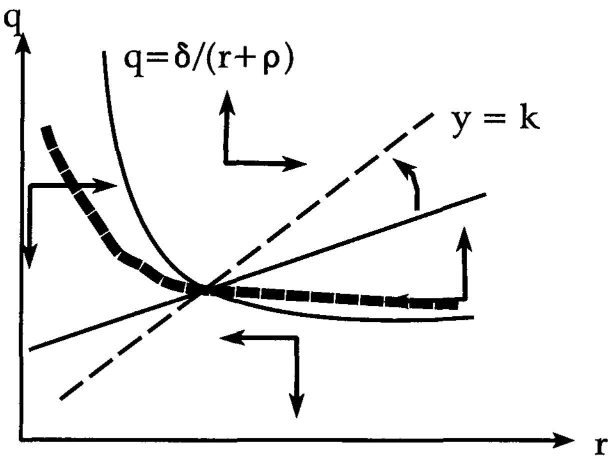
equity prices. However, levels of variables do matter because the change in equity prices can be explained by the arbitrage condition.

Indeed, the effect of this emendation of the rule is to change the extent of the central bank's feedback on both the level of equity prices and its policy instrument. While there is no change in either the slope or the position of the equity arbitrage condition, the slope of the condition at which policy rests (evaluated at the steady state) becomes:

$$\frac{\Delta(q-\bar{q})}{\Delta(r-\bar{r})}=\frac{\lambda-\alpha a_2}{\alpha a_3+\frac{\lambda\delta}{\bar{q}^2}}$$

Notice that a positive term, the feedback to equity price changes, enters the numerator. As suggested by Figure 7, increasing the responsiveness to changes in equity prices rotates the policy locus leftward. For small values of λ, this steepens the saddlepath; as a result, policy deviations produce larger swings in equity values. For sufficiently large values of λ, the dynamics become explosive so that any policy deviation from the long-run value of the real rate eventually generates an unbounded change in equity values. Essentially, by responding to past changes in stock prices, the central bank has become a feedback trader, amplifying swings in prices. Should that feedback be sufficiently intense, that behavior becomes destabilizing.

Figure 7



The message from this model for an analyst seeking to raise the weight of asset prices in monetary policy is discouraging, in that

- responsiveness to the level of equity prices increases the impact of news on valuations, while
- responsiveness to the change in equity prices sets up a feedback loop that raises the net swing in equity prices to policy misalignments and may, depending on parameters, destabilize.

If a central bank steers clear of these problems by not giving equity prices an explicit role in its policy setting, it will still have to monitor the stock market, even beyond any concerns about systemic risk. Equity prices will enter into policy formulation to the extent that they are important determinants of aggregate demand and a contemporaneous signal of the stance of policy.

Conclusion

Because of the lags in the effects of monetary policy on spending and, even further delayed, on price formation, waiting to tighten until inflation is evident is waiting too long. As a result, policy must be preemptive to help to preserve an economic expansion. However, conducting monetary policy in a preemptive fashion requires being willing to set policy on the basis of forecasts of real economic activity and inflation. One element that is important in such an assessment is the current level and the expected future direction of equity prices. Thus, elevated equity prices might tip the balance in favor of policy restraint if a forward-looking central bank was concerned that this higher wealth would fuel excessive aggregate demand or otherwise signalled that policy was misaligned. In that formulation, equity prices are important for what they imply about future spending and inflation, not for their own sake. Indeed, they have the same status as any of the other critical determinants of spending, production, and inflation, such as the foreign exchange value of the currency, the federal budget balance, and foreign economic activity.

Some analysts have suggested that equity prices, at times, should have a more important role than solely what they imply for the economic forecast, particularly when there are signs that the market is overvalued. Obviously, this requires that the central bank have a firmer view of fundamentals than the market and a willingness both to stray from its fundamental goal of achieving price stability and to accept responsibility for the actions it might take to move the market back to fundamentals. In principle, equity prices might receive elevated consideration if the central bank were worried that the inflating of an asset bubble implied a significant misallocation of resources or that the bursting of the bubble would pose systemic risks or other macroeconomic spillovers that could not be addressed by other instruments of monetary policy, including the discount window, or, perhaps, by supervisory action. The message from a simple theoretical model of the economy is that such proposals should be treated warily. While equity prices do indeed signal the stance of monetary policy, responding to them in a mechanistic manner may actually raise their volatility and, in the extreme, destabilize.

Of course, the central bank may be concerned with distortions induced by overvalued share prices. The misallocation of resources associated with elevated equity prices could take the form of excessive consumption because households spend from their higher wealth or excessive capital spending because firms find it easy to raise funds. While such stimulus to consumption raises the risk that aggregate spending would outstrip potential output, there is an offset. Imports appear also to be sensitive to equity prices, and more so proportionally than total consumption. As for capital spending, there is evidence that elevated share values contributes to the strength of that component of aggregate demand. If a boom in spending also carries with it the risk of a subsequent bust, the associated spending has fueled impressive gains in capacity, which may help to account for the recent good performance of inflation.

Appendix: The time-series properties of the variables

Typically, researchers look for a specific form of nonstationary behavior; for instance, does a series or composite of series have a unit root? The standard tests include the Dickey-Fuller (D.F.) and augmented Dickey-Fuller (A.D.F.) statistics. A complication arises because the form and distribution of any of these statistics depend on the exact form that the null hypothesis of a unit root takes. For example, the null hypothesis of a unit root may or may not have a drift term or, as emphasized by Perron (1989), structural breaks of a variety of forms. This appendix presents a collection of test statistics for the raw variables. Tests using data underlying the monthly models were estimated using observations from January 1980 to December 1996. Additionally, statistics for combinations of those variables relevant for the model

Table A-1

Is there one unit root?

Estimated using monthly data from January 1980 to December 1996

	Assuming no drift		Drift		Assuming break	
	D.F.	A.D.F.	D.F.	A.D.F.	D.F.	A.D.F.
A. Macroeconomic variables	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
1. Unemployment rate	-0.76	-1.62	-1.79	-2.45	-1.74	-2.38
2. Federal funds rate	-1.87	-1.98	-2.64	-2.99	-2.72	-3.09
Treasury yields						
3. Three-month	-1.85	-1.95	-2.58	-2.93	-2.64	-2.99
4. One-year	-1.59	-1.64	-2.45	-2.68	-2.50	-2.72
5. Three-year	-1.25	-1.39	-2.47	-2.86	-2.48	-2.87
6. Ten-year	-0.94	-1.23	-2.52	-3.07	-2.49	-3.06
7. S&P 500 earnings-price yield	-1.85	-1.96	-1.81	-2.45	-1.67	-2.36
Private yields						
8. Commercial paper (six-month)	-1.91	-1.90	-2.75	-2.91	-2.84	-3.01
9. AAA-rated Utility	-0.55	-0.94	-2.46	-3.10	-2.39	-3.07
10. Moody's A-rated	-0.81	-1.01	-2.33	-2.80	-2.30	-2.78
Inflation expectations						
11. One year ahead	-3.46	-3.25	-3.36	-2.70	-3.63	-2.82
12. Ten years ahead	-2.86	-2.36	-3.59	-2.33	-4.09	-2.55
B. Real interest rates						
13. Federal funds rate	-2.21	-2.12	-3.22	-3.34	-3.20	-3.32
Treasury yields						
14. Three-month	-2.39	-2.31	-3.23	-3.25	-3.22	-3.25
15. One-year	-2.26	-2.09	-3.21	-3.09	-3.20	-3.09
16. Three-year	-2.01	-2.14	-2.89	-3.07	-2.88	-3.06
17. Ten-year	-2.21	-2.09	-3.06	-2.95	-3.05	-2.95
18. S&P 500 earnings-price yield	-2.33	-1.93	-3.04	-2.47	-3.08	-2.54
Private yields						
19. Commercial paper (six-month)	-2.40	-2.21	-3.32	-3.21	-3.29	-3.19
20. AAA-rated Utility	-2.16	-1.99	-3.06	-2.92	-3.09	-2.94
21. Moody's A-rated	-1.89	-1.84	-2.99	-2.97	-2.99	-2.96
C. Critical values						
1%	-4.08		-4.08		-4.32	
5%	-3.47		-3.47		-3.76	
10%	-3.17		-3.17		-3.46	

Table A-2

Are there two unit roots?

Estimated using monthly data from January 1980 to December 1996

	Assuming no drift		Drift		Assuming break	
	D.F.	A.D.F.	D.F.	A.D.F.	D.F.	A.D.F.
A. Macroeconomic variables	(1)	(2)	(3)	(4)	(5)	(6)
1. Unemployment rate	-12.42	-4.38	-12.56	-4.45	-12.91	-4.66
2. Federal funds rate	-9.57	-9.09	-9.56	-9.08	-9.57	-9.09
Treasury yields						
3. Three-month	-9.93	-8.06	-9.93	-8.06	-9.94	-8.08
4. One-year	-9.96	-8.15	-9.96	-8.15	-9.97	-8.16
5. Three-year	-9.84	-7.55	-9.84	-7.55	-9.84	-7.56
6. Ten-year	-9.95	-7.07	-9.96	-7.08	-9.96	-7.11
7. S&P 500 earnings-price yield	-11.35	-5.75	-11.39	-5.78	-11.44	-5.82
Private yields						
8. Commercial paper (six-month)	-10.70	-8.77	-10.70	-8.77	-10.70	-8.77
9. AAA-rated Utility	-8.91	-6.34	-8.93	-6.36	-8.95	-6.41
10. Moody's A-rated	-9.85	-6.79	-9.85	-6.79	-9.86	-6.83
Inflation expectations						
11. One-year ahead	-19.02	-8.53	-19.12	-8.69	-19.12	-8.69
12. Ten-years ahead	-17.49	-10.56	-17.52	-10.65	-17.52	-10.64
B. Real interest rates						
13. Federal funds rate	-11.13	-9.26	-11.13	-9.28	-11.13	-9.29
Treasury yields						
14. Three-month	-12.43	-8.40	-12.44	-8.42	-12.45	-8.45
15. One-year	-13.29	-8.40	-13.30	-8.43	-13.31	-8.44
16. Three-year	-12.15	-7.67	-12.18	-7.71	-12.18	-7.72
17. Ten-year	-13.88	-7.32	-13.93	-7.39	-13.93	-7.40
18. S&P 500 earnings-price yield	-16.48	-7.41	-16.48	-7.41	-16.52	-7.47
Private yields						
19. Commercial paper (six-month)	-12.64	-8.87	-12.64	-8.89	-12.64	-8.89
20. AAA-rated Utility	-14.38	-7.61	-14.47	-7.75	-14.48	-7.77
21. Moody's A-rated	-13.70	-7.62	-13.75	-7.70	-13.76	-7.73
C. Critical values						
1%	-4.08		-4.08		-4.32	
5%	-3.47		-3.47		-3.76	
10%	-3.17		-3.17		-3.46	

– real interest rates – were calculated. Table A-1 tests for the presence of one unit root while Table A-2 checks for two unit roots. The first two columns of each table record statistics based on the assumption that there is no drift, the middle two columns posit a significant drift, while the last two permit drift and a single permanent shift in the intercept term commencing in October 1987. The Dickey-Fuller and augmented Dickey-Fuller test statistics must be negative and larger in absolute value than the corresponding critical values, which can be found in Guilkey and Schmidt (1989) and Perron (1989).

As is clear from the entries in the first table, the raw variables have at least one unit root. Further, the simple transformations involved in calculating real rates and equity-index ratios are insufficient to render those variables stationary. In other words, it will take more work to yield cointegrating relationships, and the estimation technique will have to reflect these properties. As is evident in Table A-2, though, each variable appears to have no more than a single unit root.

Table A-3 applies this same battery of tests to the quarterly data on the logarithms of spending, real equity prices (deflated by the appropriate chain-weighted price index for each category of spending), and the slope of the term structure, sampled quarterly from 1973 to 1996. For the Perron-style test of a unit root in the presence of a segmented trend, I arbitrarily assumed a single break in the fourth quarter of 1984. Again, as is evident in panel A of the table, none of the levels of these variables are stationary, but, as in panel B, differences are. Thus, these variables appear individually to have single unit roots. It is the aim of the main text to investigate how they combine to form cointegrating relationships.

Table A-3

Are there one or two unit roots?

Estimated using quarterly data from 1973:1 to 1996:4

	Assuming no drift		**Drift**		**Assuming break**	
	D.F.	**A.D.F.**	**D.F.**	**A.D.F.**	**D.F.**	**A.D.F.**
A. Is there one unit root?	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Quantities (chain-weighted)						
1. GDP	-0.26	-0.24	-1.86	-3.26	-2.28	-3.66
2. Consumption	-0.19	0.04	-1.50	-3.03	-2.24	-3.46
3. Investment	0.20	-0.65	-1.48	-3.30	-1.46	-3.22
4. Imports	0.85	0.21	-2.16	-2.80	-2.41	-3.02
5. Slope of the Treasury term structure (10-year less 3-month)	-2.71	-3.13	-2.86	-3.39	-2.86	-3.40
Real S&P 500 deflated by price index						
6. GDP	0.58	0.08	-2.46	-2.62	-2.23	-2.56
7. Consumption	0.47	-0.03	-2.47	-2.66	-2.26	-2.61
8. Investment	0.94	0.31	-2.36	-2.48	-2.07	-2.39
9. Imports	0.51	-0.43	-2.86	-2.92	-2.42	-2.85
B. Are there two unit roots?						
Quantities (chain-weighted)						
1. GDP	-6.98	-4.21	-7.00	-4.24	-7.01	-4.28
2. Consumption	-7.72	-3.86	-7.79	-3.90	-7.84	-3.94
3. Investment	-5.47	-3.80	-5.45	-3.77	-5.46	-3.77
4. Imports	-7.31	-4.21	-7.32	-4.21	-7.32	-4.22
5. Slope of the Treasury term structure (10-year less 3-month)	-8.70	-5.38	-8.67	-5.33	-8.75	-5.49
Real S&P 500 deflated by price index						
6. GDP	-7.20	-4.21	-7.49	-4.60	-7.57	-4.67
7. Consumption	-7.23	-4.26	-7.49	-4.61	-7.56	-4.68
8. Investment	-6.91	-4.08	-7.24	-4.51	-7.31	-4.57
9. Imports	-6.12	-3.83	-6.44	-4.21	-6.47	-4.24
C. Critical Values						
1%	-4.08		-4.08		-4.32	
5%	-3.47		-3.47		-3.76	
10%	-3.17		-3.17		-3.46	

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