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Introduction

This conference volume contains the papers presented at the third meeting of central bank model builders and econometricians held at the BIS on 19th-20th February 1988. It also includes the comments made by the discussants.

The topics discussed at the meeting can broadly be divided into four categories. The first session dealt with the use of VARs and SVARs in policy modelling. The paper by *Christine Gartner and Gert Wehinger* provides estimates of core inflation in nine European countries using the SVAR approach proposed by Quah and Vahey. The authors first estimate a bivariate SVAR for inflation and output growth, identifying it using the Blanchard-Quah restriction that demand shocks do not have permanent effects on output. They then define core inflation as that part of inflation which is due to demand shocks. They also extend the Quah and Vahey analysis by including changes in the nominal short-term interest rate in the SVAR. One finding is that the different measures of core inflation tend to be quite similar in most countries.

The paper by *Raf Wouters* analyses the effects of a reduction in government spending on economic activity and prices. The paper consists of three parts. In the first part the author reviews the literature on the channels through which fiscal consolidation may affect economic activity. In the second part a general equilibrium model with sticky prices is presented in which the effects of a spending cut are simulated. The sensitivity of the results to various assumptions is analysed. Finally, in the third part the author estimates a small-scale SVAR in real GDP growth, inflation and government consumption and analyses empirically the effects of a cut in government consumption on output and inflation in Belgium, Denmark and Ireland.

The final paper in this session, by *Charles Evans and Kenneth Kuttner*, is a response to recent papers which have criticised the implicit monetary policy reaction functions and the associated monetary policy shocks in VAR models. Glenn Rudebusch, for example, pointed out that Fed funds futures market provides a ready benchmark for evaluating such VAR models. He showed that forecast errors from VARs are not highly correlated with futures market surprises and that their standard deviation is often larger than that in the futures market. The authors argue that the problems of VARs are less severe than they appear, and that only minor modifications need to be made to standard VAR specifications. They argue that time aggregation makes it hard to compare the Fed funds futures surprises with forecast errors from econometric models like VARs. Secondly, they show that the correlation between shock measures is a poor measure of model forecast performance. Thirdly, they argue that reducing lag lengths and estimating over shorter samples can improve VAR forecasts.

The topics discussed in the second session of the meeting focused on modelling inflation expectations and the credibility of monetary policy. *Dinah Maclean* explores the introduction of credibility effects in the Bank of Canada's QPM model and the implications for its dynamic properties. The author starts with a brief overview of the likely effects of increased credibility on the response of the economy to various shocks and with a description of the relevant parts of QPM that implicitly incorporate credibility effects such as the calibrated sacrifice and benefit ratios, the expectations formation and the incorporation of expectations in prices. The author shows that introducing increased credibility by merely shifting the weight from the backward-looking component to the model-consistent component in the formation of price expectations does not give the expected results in response to a demand shock. This motivates the author to model credibility effects by introducing a perceived target into the formation of price expectations, the perceived target being a function of longer-term (four to five years ahead) model-consistent expectations. This is preferred over the introduction of the actual inflation target because of its flexibility and in order to avoid the big (and unlikely) announcement effects in the latter case. The chief drawback is that it is rather an ad hoc way of introducing credibility effects.

Antulio Bomfim and Flint Brayton take a somewhat different route in their paper by estimating two types of "learning" model to capture the formation of private sector expectations about the inflation target. They assume private agents derive the central bank's unobserved inflation target from the estimation of a reduced-form policy reaction function based in one case on rolling regressions and using in the other the Kalman filter. They then compare these learning models with actual series of long-run inflation expectations derived from surveys and find that the behaviour of the estimated and survey series is comparable. In the second part of their paper, the authors use dynamic simulations of the FRB/US model to estimate the cost of disinflation under different specifications of private sector's "learning" of the new policy target. They generally find that the sacrifice ratio under these learning rules is larger than consensus estimates suggest.

Hans Dillén and Elisabeth Hopkins focus on the term structure of interest rates as a source of information concerning inflation expectations. They argue that there are two major explanations why Swedish forward interest rates have been high and volatile. First, investors' fears that the economy will switch to a high-inflation regime have given rise to a regime shift premium. Secondly, expectations of monetary policy actions have amplified the effect on forward interest rates originating from fluctuations in inflation expectations. In an empirical investigation the quantitative importance of adjusting forward interest rates for regime shift premia is shown. The authors also show that an increase of the one-year forward interest rate (adjusted for the regime shift premium) is only partially reflected in an increase in investors' inflation expectations (obtained from surveys). According to the authors this suggests that the rest of the forward interest rate movement reflects expectations of future increases of the real short-term interest rate. Finally, the authors present evidence that investors' expectations obtained from surveys also partly reflect regime shift expectations.

The third session dealt with the evaluation of different targeting regimes and use of a monetary conditions index. *Paul Conway, Aaron Drew, Ben Hunt and Alasdair Scott* perform stochastic simulations using the FPS model of the Reserve Bank of New Zealand to assess which price index to target. In particular, they compare targeting CPI inflation with targeting domestic goods price inflation. The authors find that targeting domestic price inflation reduces the variance in real output, nominal interest rates, the real exchange rate and domestic price inflation with very little increase in CPI inflation variability. The result appears to be robust even if the monetary authority is uncertain about the true expectations process and even if direct exchange rate effects influence agents' inflation expectations. Tracing out the efficient output/CPI inflation variability frontiers under both CPI and domestic price inflation targeting illustrates that the result is not limited to the base-case PFS reaction function.

Günter Coenen compares inflation and monetary targeting using a small stylised theoretical P-star model. On the basis of a deterministic simulation of the effects of a real aggregate demand and a money demand (velocity) shock under both regimes, the author concludes that monetary targeting should be the preference of a central bank whose goal is to control inflation.

The last paper in this session, by *Neil Ericsson, Eilev Jansen, Neva Kerbeshian and Ragnar Nymoen*, analyses the use of a monetary conditions index (MCI) as an operational target for monetary policy. The authors describe and define the concept, summarise how central banks implement MCIs in practice, review some of the operational and conceptual issues involved and evaluate the sensitivity of MCIs to an inherent source of uncertainty in their calculation. On the basis of this analysis, they conclude that uncertainty typically renders MCIs uninformative for their ostensible purposes and briefly consider some possible alternatives.

The fourth and final session of the meeting contained two papers on the credit channel of the monetary transmission mechanism. *Ignacio Hernando* gives an extensive overview of the literature on the credit channel and examines the response of the financing "mix" (the ratio of bank loans to commercial paper) in the Spanish business sector to changes in the stance of monetary policy. He finds weak evidence that the mix changes away from bank loans after a policy tightening, and stronger evidence showing a widening of the yield spread between loans and commercial paper.

Olivier Steudler and Mathias Zurlinden, on the other hand, focus on the effects of monetary policy on the composition of the balance sheets of Swiss banks. Following the work of Kashyap and Stein, they analyse the differences in the response of loans, deposits and bonds issued for three groups of banks (the big three, the cantonal and the regional banks) to a monetary policy tightening. The Kashyap and Stein hypothesis predicts that such a tightening causes a relatively strong reduction of small banks' loan portfolio and of large banks' securities holdings. Based on the point estimates of the impulse response functions, the authors find that the responses are consistent with the predictions for securities but not for loans. In contrast to the hypothesis, the three big banks seem to experience the sharpest decline in loans after an interest rate shock.

Participants in the meeting

Australia:	Mr. Gordon DE BROUWER
Austria:	Mr. Gert WEHINGER Ms. Christine GARTNER
Belgium:	Mr. Philippe MOËS Mr. Raf WOUTERS
Canada:	Ms. Dinah MACLEAN Mr. David TESSIER
Finland:	Ms. Hanna-Leena MÄNNISTÖ
France:	Mr. Pierre SICSIC Mr. Franck SÉDILLOT
Germany:	Mr. Wilfried JAHNKE Mr. Günter COENEN
Italy:	Mr. Filippo ALTISSIMO Mr. Francesco LIPPI
Japan:	Mr. Koichiro KAMADA
Netherlands:	Mr. Peter van ELS Mr. Carsten FOLKERTSMA
New Zealand:	Mr. Ben HUNT Mr. Aaron DREW
Norway:	Mr. Eilev JANSEN
Spain:	Mr. Ignacio HERNANDO
Sweden:	Mr. Hans DILLÉN Ms. Elisabeth HOPKINS
Switzerland:	Mr. Mathias ZURLINDEN Ms. Barbara LÜSCHER
United Kingdom:	Mr. Alec CHRYSTAL
United States:	Mr. Kenneth KUTTNER *(New York)* Mr. Antulio BOMFIM *(Washington)* Mr. Chris ERCEG *(Washington)* Mr. Neil ERICSSON *(Washington)*
BIS:	Mr. Renato FILOSA (Chairman) Mr. Joseph BISIGNANO Mr. Zenta NAKAJIMA Mr. Palle ANDERSEN Mr. Claudio BORIO Mr. Gabriele GALATI Mr. Stefan GERLACH Mr. Robert McCAULEY Mr. Frank SMETS Mr. Kostantinos TSATSARONIS

Core inflation in selected European Union countries

Christine Gartner and Gert D. Wehinger*

Introduction

The issue of how to measure inflation and, in particular, its underlying trend has attracted increasing attention in recent years. A major reason for this renewed interest is that a number of central banks, both inside and outside the European Union, have committed themselves to explicit quantitative inflation targets.[1] The assessment of deviations of current and expected inflation from the target requires taking volatile and temporary price influences into account. The issue of distinguishing transitory from persistent price movements is also relevant for countries aiming for price stability in other monetary policy frameworks than inflation targeting. Alternative inflation indicators, especially those of underlying inflation, may cast light on the sustainability of a country's inflation performance.

An important limitation of commonly used inflation measures such as the Consumer Price Index (CPI) is their susceptibility to specific disturbances which are unrelated to the "pure" (or core) inflationary process. As a result, measured inflation may give a misleading picture of underlying price trends relevant for monetary policy.

The purpose of this study is to provide information on underlying price movements relevant for the single monetary policy of the ECB. For comparative reasons, we use a model-based approach to calculate core inflation indicators for selected European countries. The core inflation process is identified by means of a VAR (vector autoregression) technique that was first suggested by Quah and Vahey (1995). We use a modification of the original model along the lines specified by Blix (1995) and Dewachter and Lustig (1997) in order to split measured inflation into core and non-core components. The underlying inflation process is that component of measured price movements which is governed by demand shocks.

In view of the central role price stability plays for the single monetary policy of the ECB alternative inflation indicators, especially those of core or underlying inflation, will play an important role as monetary policy indicators, independent of the specific choice of the monetary policy strategy by the ECB. Although this topic has been treated in some studies, Austria has never been included so far.

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1 For a comprehensive survey see, for example, Leiderman and Svensson (1995), and Haldane (1995). More recent contributions include Debelle (1997) and Masson et al. (1997).

1. The concept and measurement of underlying inflation

Although the concept of underlying inflation is widely used in monetary policy analysis,[2] views differ about its precise definition.

Most papers[3] refer to Eckstein's (1981) definition of underlying or core inflation as the rate of price increases that would occur along the economy's long-term growth path. The core inflation rate is thus a steady-state concept and equivalent to the *trend increase of the price of aggregate supply.* Alternatively, Parkin (1984) assumes that in the long-run equilibrium, factor prices for labour and capital fully reflect inflation expectations. In that case, core inflation is identical to *expected inflation.* As deviations of actual from core inflation result from demand fluctuations and random supply (and other) disturbances, his results are consistent with the existence of a short-run expectations-augmented supply (or Phillips) curve reflecting such factors.

As there is no single concept of what is understood by core inflation it is not surprising that views on how to measure it differ.

The standard approach has been to remove, in some ad hoc manner, the "unwanted" component, such as transitory noise, which has its sources in changing seasonal patterns, resource shocks, exchange rate changes, indirect tax changes or asynchronous price adjustments, or other distorting influences like weighting differences, quality changes, new goods or the substitution bias. The remainder is seen as a reliable estimate of the underlying inflation process.[4] Removing distorting, temporary or particularly volatile influences can be done either on a case-by-case basis or in a more structured way. The first group of procedures includes the zero-weighting technique and its variants. The structural methods of calculating specific underlying inflation indicators include simple as well as more sophisticated smoothing techniques (trimmed mean method; Hodrick-Prescott filter, Kalman filter) and the VAR models based on the paper by Quah and Vahey (1995). Model-based calculation of core inflation allows an economic interpretation of the resulting indicator. In contrast, in the case of ad hoc procedures such as zero-weighting and smoothing techniques, an interpretation based on economic theory is not straightforward.

We decided to use a VAR approach similar to Quah and Vahey's for two reasons:

1. Fluch and Gartner (1997) suggest that mechanical procedures such as the zero-weighting approach have certain drawbacks for cross-country analysis. Their empirical results show that the trend of and deviations from headline inflation heavily depend on the definition used. In spite of harmonisation efforts initiated by the European Monetary Institute, concepts of calculating core inflation still differ markedly.
2. We are interested in a forward-looking assessment of inflation performance. Forecasting is not possible with the zero-weighting procedure and possible only with certain restrictions using the smoothing technique, whereas a model-based approach enables to project historical structures into the future.

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2 The interest in Austria in alternative inflation indicators is relatively new. As is well known, the Oesterreichische Nationalbank (OeNB) follows an exchange rate target and thus gears its monetary policy to that of the anchor currency (among others, see Gartner (1995), and Hochreiter and Winckler (1995)). The effectiveness of the monetary strategy is measured in terms of the degree of inflation convergence with Germany. Up to now measures of underlying inflation played only a limited role. As far as the OeNB is concerned it focused its attention on the headline inflation rate, the CPI changes being the inflation indicator, making additive adjustments for the contribution of specific indirect tax changes or seasonal food prices whenever relevant.

3 Among others, see EMI (1995).

4 This approach has been used, inter alia, in Sweden, the United Kingdom and Finland, and was also suggested by the EMI.

2. Identifying core inflation

The two approaches mentioned above (zero weighting and smoothing) remove, in some *ad hoc* manner, the "unwanted" components ("noise") of measured inflation. What remains ought to be a reliable estimate of the underlying inflation process. In their paper, Quah and Vahey (1995) argue that the conceptual mismatch between current methods for calculating inflation and economic theory is more than just a measurement error. Price indices such as the CPI measure the costs of particular goods and services, while the economic notion of inflation is that of sustained increases in the general price level. As economic theory does not suggest a particular functional form of inflation, there is no justification for believing that core inflation is the result of some arbitrary smoothing procedure.

Consequently Quah and Vahey (1995) suggest an alternative technique that is based on an explicit economic hypothesis. They define core inflation as that component of measured inflation that has no medium to long-run impact on real output. This definition is consistent with a vertical long-run Phillips curve interpretation of the co-movements in output and inflation. They then implement this definition as a restriction on a bivariate SVAR (structural vector autoregressive) model and use it to extract a measure of core inflation. Our identification scheme differs only slightly as we identify effects on prices instead of price changes, thus referring to, from a theoretical viewpoint, a standard aggregate demand/aggregate supply framework.[5]

2.1 Methodology

The identification scheme of Quah and Vahey's model is very similar to that of Blanchard and Quah (1989) and Shapiro and Watson (1988).

It follows the VAR tradition in methodology, employing impulse response analysis and variance decompositions. The identification of the shocks is based on a Choleski decomposition of a long-run parameter matrix and is therefore different from the short-run identification schemes of Bernanke (1986) and others.

The structural model of real GDP, y, and CPI, p, has the long-run solution form:

$$y = f(\varepsilon^s) \text{ and } \tag{1}$$

$$p = f(\varepsilon^s, \varepsilon^d) \tag{2}$$

We assume that the economy is hit by innovations given in the vector $\varepsilon = (\varepsilon^s, \varepsilon^d)$, which consists of a supply shock ε^s and a demand shock ε^d. While supply shocks[6] may have permanent effects on both prices and output, demand shocks are defined to have no long-run effect on output, i.e. they are transitory with respect to real variables. We identify the core inflation process as that part of the increases in the CPI that has no long-run effects on output, i.e. price movements that are determined solely by shifts in the aggregate demand curve ("demand pull" inflation).[7] We compute core inflation by simulations imposing paths of structural shocks as described in Section 2.3.

We impose two kinds of restrictions on structural innovations. First, both of the structural disturbances are assumed to be uncorrelated at all leads and lags and have unit variance. Second, demand shocks cannot have long-run effects on output. The long-run effects of demand disturbances

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5 From an empirical viewpoint we refer to the fact that most price changes can be considered as (trend-)stationary. See also the data section below on this issue.

6 Typical supply shocks are productivity changes, energy shocks, taxes and price controls.

7 The simple framework applied here could be extended in order to capture also, for example, "cost-push" inflation effects by including other variables such as wages and other specific prices.

on CPI are unconstrained. These restrictions are sufficient to uniquely identify both of the underlying disturbances as will be shown below.

2.2 Identifying restrictions and identification of the model

Assume that a vector Δx of (differenced) macroeconomic variables follows a covariance stationary process of the form:

$$\Delta x_t = \mathrm{C}(L)u_t \tag{3}$$

In our case $\Delta x = [\Delta y, \Delta p]'$, with y the log of domestic output and p the log of prices (CPI), respectively. C(L) is a lag polynomial where the C's are coefficient matrices at the respective lags of the serially uncorrelated errors u, where $E(u\ u')=\Sigma$. The first coefficient matrix of the polynomial, C_0, is normalised to the identity matrix I.

A reduced form and normalised moving average representation of the same process is given by:

$$\Delta x_t = \mathrm{E}(L)e_t \tag{4}$$

with $\mathrm{E}(e\ e') = I$ and the shocks uncorrelated across time and across variables.

Only the u's can be directly estimated from the VAR, the e's have to be calculated based on its moving average representation (3). As we have assumed $C_0 = I$ and we have a linear relation between C(L) and E(L) we can write:

$$u_t = E_0 e_t \tag{5}$$

The problem is then to find E_0 imposing $k \times k$ restrictions, where k is the number of variables in the model and thus $k \times k$ is the dimension of E_0.

From $ee' = I$ and $uu' = \Sigma$ we have with (5):

$$\Sigma = E_0 E_0' \tag{6}$$

This factorisation yields $\frac{1}{2}k(k+1)$ non-linear restrictions, for the rest of $\frac{1}{2}k(k-1)$ restrictions we impose long-term neutrality properties for certain errors driving the respective variables. If we evaluate the polynomial matrices at L = 1, where a matrix $\mathrm{E}(1) = E_0+E_1+E_2+E_3...$, we get the long-run impacts of errors on the variable vector Δx, and, specifically,

$$\Delta^* x = \begin{bmatrix} \Delta^* y \\ \Delta^* p \end{bmatrix} = \begin{bmatrix} E_{11}(1) & 0 \\ E_{21}(1) & E_{22}(1) \end{bmatrix} \begin{bmatrix} \varepsilon_s \\ \varepsilon_d \end{bmatrix} \tag{7}$$

where $\Delta^* x = \lim_{t \to \infty} x_t - x_{t^*}$

As E(1) is assumed to be lower triangular, we can use this fact to recover E_0 in the following way. Equating (3) and (4) at their long-run values we have

$$\mathrm{C}(1)u_t = \mathrm{E}(1)e_t. \tag{8}$$

With $ee' = I$ and $uu' = \Sigma$, the matrix E(1) can be derived from a Choleski decomposition of

$$\mathrm{C}(1)\Sigma \mathrm{C}(1)' = \mathrm{E}(1)\mathrm{E}(1)' \tag{9}$$

From the values for C(1), which can be derived from the estimated VAR-parameters, and the variance-covariance matrix Σ we compute the Choleski factor E(1) and can then recover E_0 as:

$$E_0 = \mathrm{C}(1)^{-1}\,\mathrm{E}(1) \tag{10}$$

The matrix E_0 can then be used in $u_t = E_0 e_t$ to compute the impact of structural shocks on the elements of Δx_t (orthogonal impulse responses).

With this background, we proceed as follows for the empirical analysis. First we estimate a vector-autoregressive (VAR) model of the form:

$$\mathrm{A}(L)\Delta x_t = u_t \tag{11}$$

From A(L) we compute (accumulate for) the long-run entries of A(1). Inverting, yields $\mathrm{A}(1)^{-1} = \mathrm{C}(1)$. Consequently we get E_0 from (9) and (10), which we use to compute the respective impulse responses and the variance decomposition of the structural shocks given in (4).

2.3 Computing core inflation: simulations using structural shocks

We calculate core inflation by imposing certain paths of structural shocks. The structural shocks e_t are recovered from the estimated errors u_t through the relation $e_t = E_0^{-1} u_t$. Having found e_t two alternative forecast simulations can be computed by dropping certain elements of the shock vector: the variables' path "due to" specific, single shocks and "absent" specific shocks.

The first class of simulations can be done by setting $\mathrm{e}_t^S = [\mathrm{e}_{S,t}, 0]$ for the simulations "due to supply" and $e_t^D = [0, e_{D,t}]$ for the simulations "due to demand", where the errors u_t^X (X=S,D) to be used for the forecasts with the estimated VAR models will be recovered from $u_t^X = E_0 e_t^X$.

The alternative simulations pursued here set $\mathrm{e}_t^{S'} = [0, \mathrm{e}_{D,t}]$ for the simulations "absent supply" and $\mathrm{e}_t^{D'} = [\mathrm{e}_{S,t}, 0]$ for the simulations "absent demand", where, as before, the errors $\mathrm{u}_t^{X'}$ $(X' = S', D')$ to be used for the forecasts with the estimated VAR models will be recovered through $u_t^{X'} = E_0 e_t^{X'}$.

As the originally estimated variables are differences, we also perform accumulations (eventually including a mean that had been subtracted before estimation) in order to see how the simulated levels of the variables would evolve under the different assumptions.

Core inflation π^c is defined as that component of inflation which has no permanent effect on output. In our specification that would correspond to the "absent supply" or the "due to demand"[8] simulation path for Δp.

2.4 Interpretation

The first important assumption underlying this technique concerns the number of structural innovations. Quah and Vahey (1995) assume that there are only two types of shocks affecting inflation and output. In reality, the economy is hit by a large number of heterogeneous shocks, and each of them may have different effects on measured inflation and output. In line with the

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8 This is only true for the bivariate SVAR system, of course.

work of Blix (1995) and Dewachter and Lustig (1997), we explicitly address this potential misspecification problem by extending the SVAR and checking the robustness of the results. In the extension we distinguish between monetary and real aggregate demand shifts, since these may affect inflation and output differently.

The second debatable assumption is the orthogonality restriction on the structural innovations. Following the Quah and Vahey (1995) methodology we assume core and non-core innovations to be uncorrelated at all leads and lags. Nevertheless, some policy shifts in response to core shocks (for instance a restrictive or loose fiscal policy in response to a price hike) may have a permanent effect on output. As a result, non-core innovations may be caused by core innovations. The model, however, excludes the possibility of actual correlation.

The identifying restrictions do not constrain the structural multipliers determining the response of measured inflation to non-core innovations. This long-run effect is entirely determined by the estimations. If these non-core innovations explain a sizeable part of the long-run variability in measured inflation, the Quah and Vahey (1995) identification procedure has to be re-examined. This would mean that the non-core innovations drive the underlying inflationary process.

2.5 Extension: including monetary policy

To assess the restrictiveness of the two-shock approach outlined above, we extend the bivariate SVAR by introducing a monetary variable. This has been done before: Blix (1995) introduced monetary aggregates as a third variable. Dewachter and Lustig (1997), who are mainly interested in empirical results for the ERM-countries, include a short-term nominal interest rate in the model. As our (future) interest is in common trends in underlying inflation, we proceed along the lines of Dewachter and Lustig (1997) and also include short-term interest rates as the monetary policy variable. We implicitly assume that monetary aggregates are endogenous, which appears to be a fair assumption for most European countries.

We assume that a small open economy with a fixed exchange rate regime is hit by three structural innovations: a supply shock, a monetary shock and a demand shock, the latter two of which are core innovations. Hence, the structural model in real output, y, short-term interest rates, i, and CPI, p, in its long-run representation has the following form:

$$y = f(\varepsilon^s), \tag{12}$$

$$i = f(\varepsilon^s, \varepsilon^m), \text{ and} \tag{13}$$

$$p = f(\varepsilon^s, \varepsilon^m, \varepsilon^d) \tag{14}$$

The non-core innovations ε^s are interpreted as supply disturbances (e.g. technology shocks),[9] which generate relative price shifts. These supply shocks are assumed to have a permanent effect on output. As before core inflation is defined as that component of measured inflation which is not affected by supply innovations.

The first type of core innovations ε^m captures the effects of a monetary disturbance. These LM-innovations do not affect real output permanently, but they are supposed to exert a lasting influence on short-term nominal interest rates and on inflation. Given the validity of interest parity, $i = i^* + \dot{e}$,[10] in the long run, the ε^m innovation can also be interpreted as an EU-wide (ERM-wide, see

9 Cf. footnote 6.

10 Where i denotes the domestic interest rate, i^* the foreign interest rate or that of the anchor currency country and $\dot{e}$ is the expected change in the nominal exchange rate over time.

below) monetary policy shock. As for countries pursuing a fixed exchange rate regime it holds that $\dot{e} \cong 0$ in the long run, an exogenous shift in the level of i^* has to be accommodated by a permanent shift in i. In the short run, due to lower credibility of the peg, i can deviate from i^* to the extent of devaluation expectations.

Two major effects of nominal interest rate innovations can then be distinguished among countries of the European Monetary System (EMS): for (smaller) countries with a credible and tight exchange-rate peg (within the Exchange Rate Mechanism, ERM) an interest rate increase will arise mainly due to an accommodation of an increased ERM-wide interest rate level, and even short-run output and price effects should be very small. For countries allowing (or having allowed) for more flexibility in the exchange-rate peg (e.g., not having permanently participated in the ERM) a nominal interest-rate shock can, given the validity of the interest parity, also be interpreted as following an autonomous expansionary monetary disturbance, giving rise to devaluation expectations $\dot{e}$, increasing output at least temporarily (long-run effects are restricted to be zero) and prices even at longer time horizons.[11]

The second type of core innovations consists of a real demand shock. This AD- or IS-shift affects the rate of inflation in the short run and the price level in the long run, but leaves output and the interest rate level (i) unchanged at an infinite horizon.

Consider a vector Δz which now includes changes in the short-term nominal interest rate, Δi. This vector Δz is a covariance-stationary process not constrained by a cointegrating relation. This in turn means that it has an invertible moving average representation which, in its long-run (accumulated) form, is given by:

$$\Delta^* z = \begin{bmatrix} \Delta^* y \\ \Delta^* i \\ \Delta^* p \end{bmatrix} = \begin{bmatrix} E_{11}(1) & 0 & 0 \\ E_{21}(1) & E_{22}(1) & 0 \\ E_{31}(1) & E_{32}(1) & E_{33}(1) \end{bmatrix} \begin{bmatrix} \varepsilon_S \\ \varepsilon_M \\ \varepsilon_D \end{bmatrix} \tag{15}$$

where $\Delta^* z = \lim_{t \to \infty} z_t - z_{t^*}$

ε_S denotes the supply shock (i.e. non-core innovation), ε_M represents the monetary shock and ε_D is a real demand disturbance. Note that the matrix of the structural multipliers in (15) is invertible. This system is fully identified. Core innovations are distinguished from non-core innovations by imposing that the latter cannot affect output in the long run. Money demand shocks are distinguished from real demand innovations by assuming that the latter have no lasting impact on interest rates.

2.6 Computing core inflation in the extended model

As in Section 2.3 we again use the method of imposing long-run paths on structural shocks to compute core inflation π^c. Having recovered e_t from the estimated errors u_t through the relation $e_t = E_0^{-1} u_t$, two alternative forecast simulations are computed by dropping certain elements of the shock vector: the variables' path "due to" specific, single shocks and "absent" specific shocks.

In the trivariate case, the first class of simulations can be done by setting $e_t^S = [e_{S,t}, 0, 0]'$ for the simulations "due to supply", $e_t^{LM} = [0, e_{LM,t}, 0]'$ for the simulations "due to

[11] In fact, as shown below, we find such behaviour of variables in Belgium, Finland, France, Italy, Sweden and the United Kingdom.

LM" and $e_t^D = [0, 0, e_{D,t}]$ for the simulations "due to demand ", where the errors u_t^Z ($Z = S, LM, D$) to be used for the forecasts with the estimated VAR models are recovered from $u_t^Z = E_0 e_t^Z$.

The alternative simulations pursued here set $e_t^{S'} = [0, e_{LM,t}, e_{D,t}]$ for the simulations "absent supply", $e_t^{LM'} = [e_{S,t}, 0, e_{D,t}]$ for the simulations "absent LM", and $e_t^{D'} = [e_{S,t}, e_{LM,t}, 0]$ for the simulations "absent demand", where, as before, the errors $u_t^{Z'}$ ($Z' = S', LM', D'$) to be used for the forecasts with the estimated VAR models are recovered through $u_t^{Z'} = E_0 e_t^{Z'}$.

As the originally estimated variables are differences, we perform accumulations as in the bivariate case.

Again, core inflation π^c is defined as that component of inflation which has no permanent effect on output. In the trivariate SVAR model this would correspond to the "absent supply" simulation path for Δp, as core inflation is only that component of measured inflation which is driven by core (real demand and monetary) shocks.

3. Estimation

In this section we apply the identification technique outlined above to assess the performance of the CPI as a measure of "true" inflation. This is done simply by tracing the difference between measured inflation (using CPI) and (computed) core inflation using bivariate and trivariate SVAR models. We estimate bi- and trivariate VAR systems in GDP growth, changes in prices and short-term nominal interest rates for Austria, Belgium, Germany, Finland, France, Italy, the Netherlands, Sweden and the United Kingdom. The estimation period is 1971:1 to 1996:4. Values for 1997 and 1998 are forecasts from the estimated VAR model.

3.1 Data

We use quarterly, non-seasonally-adjusted data for the CPI (or a comparable price index such as cost of living or Retail Price Index – RPI) provided by OECD Main Economic Indicators. Quarterly GDP data and short term interest rates (3-months) are taken from the BIS data base. We subject the log levels of the data to a couple of tests such as the Hylleberg test,[12] the Augmented Dickey-Fuller[13] (ADF) as well as the Phillips-Perron[14] tests. The Hylleberg test results suggest to take the fourth lag differences of the data, ADF and Phillips-Perron tests are then applied to these differences. The results are broadly consistent with output, prices and interest rates being integrated of order one (hence, there is at least one shock for each variable affecting it permanently). Therefore, GDP, prices and interest rates enter the VAR system as year-on-year growth rates. Before entering the VAR, we deduct the respective means from changes in GDP and interest rates (i.e. the level series contain a trend). As the test results suggest year-on-year inflation rates to be trend-stationary, we adjust inflation rates for a trend variable, which could capture the impact of a "secular" downward

[12] Hylleberg et al. (1990) suggest a test for seasonal roots, as implied by our annual differencing of the data.

[13] See Dickey and Fuller (1979,1981).

[14] See Perron (1988) and Pillips and Perron (1988).

trend in inflation which is observed in most countries.[15] Such a behaviour of inflation seems plausible, given the increase in competitive pressures, the ongoing deregulation and integration of markets; at least, test results in general do not suggest cointegrating restrictions or error correction terms.[16]

3.2 Bivariate SVAR

As a first step bivariate VAR systems in GDP growth and changes in prices are estimated over the period 1971:4 to 1996:4 for all countries. We include three lags, supported by various information criteria.[17] Estimation results are reported in Figures 1 to 9.[18] Both inflation measures (CPI and core inflation) are calculated as the log change in the price level with respect to the corresponding quarter of the previous year. Core inflation is estimated as specified in Section 2.3.

3.2.1 Core versus CPI inflation

Figure 1 displays the results for ***Austria***. Overall CPI inflation seems to track the underlying rate of inflation reasonably well. The peaks and troughs of both measures coincide more or less. Yet the deviations tend to be very persistent. From 1971 to 1975 the underlying inflationary process was stronger than the conventional inflation measure would have suggested. After 1975 the opposite was true. Beginning with the late 1970s up to 1987 CPI inflation was considerably higher than our measure of core inflation resulting mainly from the absence of positive supply shocks (productivity slowdown). In the late 1980s, the Austrian economy was hit by a number of positive demand (core) shocks which led to an underlying inflation process considerably stronger than CPI inflation.

Estimation results for ***Belgium*** are shown in Figure 2. Again, core inflation tracks actual inflation quite well. We found a core inflation process that is in some periods considerable weaker than actual inflation. Especially, in the years around the first (1974) and the second oil price shock (1981) inflation was overestimated by the conventional inflation statistics. Also in the 1990s core inflation is lower than actual inflation. After 1993, deviations of core from actual inflation diminish gradually due the absence of positive supply shocks. At the end of 1993, the "plan global" was implemented which included tax increases and programmes of wage moderation. Consequently, core shocks gained relative importance explaining the inflation process.

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[15] Many price series can be considered borderline cases between being I(1) and I(2) (integrated of order one or two, respectively). As we found I(1) evidence in many cases we treated even the borderline cases as such in order to provide a single framework for our analysis.

[16] Applying the Engle and Granger (1987) tests we could not find cointegrating relationships between the variables; applying Johansens (1991) procedure some of the cases look more ambiguous. However, adding error correction terms to the VAR then did not seem to alter the results significantly. Therefore and in order to keep the framework simple but still applicable to all countries we did not estimate the model in its vector-error correction form.

[17] Three information criteria were used to determine the lag length for the respective VAR estimation: the Akaike Information Criterion (AIC; Akaike (1973), the Schwarz Information Criterion (SC; Schwarz (1978); for both cf., e.g., Judge et al. (1988), p. 870ff), and the Hannan and Quinn Information Criterion (HQ; Hannan and Quinn (1979)), using, respectively, the simple formulae

$$AIC = \log|\Sigma| + \frac{2k}{T}, \quad SC = \log|\Sigma| + \frac{k \log T}{T}, \quad HQ = \log|\Sigma| + \frac{2k \log(\log T)}{T},$$

where $|\Sigma|$ is the determinant of the variance-covariance matrix of the VAR residuals, k is the number of parameters in the model and T is the number of observations.

[18] All graphs can be found in Appendix B.

Estimation results for ***Finland*** can be seen in Figure 3. According to our calculations the Finnish case represents an exception. Very much like the British RPI, the Finnish CPI inflation seems hardly to be influenced by core innovations. Supply shocks tend to have had a massive impact on the Finnish inflation statistics. Deviations of the underlying inflation measure from the CPI inflation process are substantial. Massive positive deviations can be observed for the years around the first and the second oil price shocks. More recently the opening up of Eastern Europe had significant consequences for the Finnish economy. Negative supply shocks lead to an underlying inflation rate considerably lower than actual inflation. The danger of imported inflation due to a sharp depreciation of the markka was mitigated by incomes policy. In more recent years the core inflation indicator overestimated actual inflation (which could be a sign of an overheating economy).

The inflation experience in ***France*** is illustrated in Figure 4. We find an underestimation of the underlying inflation by the conventional inflation statistics in the first part of the 1980s, whereas in the second half of the decade inflation was underestimated by the CPI measure. For the 1990s, we get a core inflation measure that is lying substantially below measured inflation. One explanation could be that the French economy, in the process of budget consolidation, was hit by a couple of supply shocks that are not captured by the core inflation measure.

Figure 5 considers the case in ***Germany***. As in the Austrian case, the calculated core inflation tracks the CPI inflation well, i.e. the turning points coincide. The deviations of core inflation from CPI are not very large; with the exception of 1991 (German unification) they remain within the 1.5% band over the whole sample.

The results for ***Italy*** are summarised in Figure 6. The Italian CPI seems to perform very well in measuring inflation. The differences between the two inflation measures are minor. There is also evidence that supply shocks have had only a very restricted impact on the CPI inflation measure. As a result the calculated underlying inflation process perfectly fits the CPI inflation.

Figure 7 shows the estimated core and CPI inflation for the ***Netherlands***. The assessment of our results for the Dutch inflation experience is very much the same as for Italy. Supply shocks seem to have only a minor impact on the inflationary process. The deviations of actual inflation from core inflation remain well within the 1% band. As for Italy, we have no clear-cut explanations for these empirical findings.

As can be seen from Figure 8, our calculations for the underlying inflation rate follow the CPI measure considerably well also in ***Sweden***. At the beginning of the sample, the underlying inflation indicator ignores the ups and downs of the rather volatile CPI inflation rate. So we cannot give a clear statement whether the underlying inflation rate was definitely over- or underestimated in the first part of the 1970s. In the second part of the 1970s core inflation is overestimated by CPI inflation. The picture changes at the beginning of the 1980s: Deviations of actual inflation from core inflation tend to be comparatively small in the 1980s due to the absence of positive supply shocks. Negative supply shocks and a strong depreciation of the krona led to an actual inflation rate that substantially overestimated the underlying inflation rate. Beginning with 1994 price stability could be restored. In the following years the Swedish economy displayed low inflation rates, hence it is not surprising that the calculated core inflation indictor is well above the measured CPI inflation.[19]

The results for the ***United Kingdom*** are reported in Figure 9. The calculated core inflation measure for the United Kingdom tends to be relatively smooth as compared to the actual inflation. This means that supply innovations seem to have an important impact on the measured inflation rate. As the United Kingdom is one of the major oil producing OECD countries (apart from Norway), oil price shocks constitute an important (and positive) part of supply shocks leading to downward shifts of the price level. Consequently, actual inflation overestimates the underlying

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[19] The results for this period are completely opposed to the observations by Blix (1995). He found a strong overestimation of the core inflation by the CPI measure. Thereafter the core inflation calculated by Blix shows a smoother development as is the case with our calculations.

inflation trend for the respective periods. In the 1980s the absence of positive supply shocks brings about an underlying inflation that lies considerably above the measured inflation rate (which could also be due to the influence of low oil prices, a non-core element of the inflation process). At the beginning of the 1990s the calculated core inflation rate is very low and turns out to be negative for a few periods. Negative productivity shocks may have pushed RPI above core inflation. Towards the end of the sample, positive productivity shocks (increased flexibility of the labour market) may have put downward pressure on inflation by increasing the output potential and thus resulting in an underlying inflation lower than the usual inflation measure.[20]

We compared our findings with those of Bjørnland (1997), Blix (1995), Dewachter and Lustig (1997), Fase and Folkertsma (1997), Quah and Vahey (1995) and Jacquinot (1998), who used similar concepts. It is not surprising that their results sometimes differ markedly. We want to name only three possible reasons for these differences, which seem to be the most influencing factors. First, in contrast to other empirical studies on this topic, we did not use industrial output data as a proxy for overall output of the economy, but we applied real GDP.[21] Due to data availability, the second difference is a consequence of the first: we used quarterly instead of monthly data. The third source for the deviation clearly comes from the specification of the model. As we assumed the inflation rate to be (trend-)stationary, the change of prices instead of the change of the inflation rate enters the VAR system. The results are very sensitive to such differences in specification.

3.2.2 Impulse response functions and variance decompositions

Figures 10, 14, 18, 22, 26, 30, 34, 38 and 42 report the estimated dynamic responses of measured inflation and output to a one percentage point (ppt.) supply (core) and demand (non-core) shock, for all countries and for the bivariate case. For our purposes the upper and lower right graphs of each figure are relevant.

The dynamic response of CPI inflation to supply disturbances differs substantially from its response to demand disturbances. The results for the impulse response functions very much coincide with what we would expect from theory. Let us consider a simple AS–AD (aggregate supply–aggregate demand) model. A positive productivity shock would shift the AS curve to the right. As a consequence, prices would decrease. This is exactly what we can see in the shape of the impulse response function of CPI on a one period one ppt. increase in aggregate supply. An initial downward jump in prices is followed by step-by-step decreases of prices until the inflation rate converges to zero and the new price level is found.

A positive demand shock shifts the AD curve to the right. In the absence of price rigidities, we would observe immediate price increases. In any case, prices adjust until the new equilibrium is reached. The adjustment process of prices gives us the shape of the impulse response function of CPI to a one ppt. increase in aggregate demand. Immediately after the demand shock an increase in the price level can be observed. After that inflation rates decrease step-by-step until the inflationary impact of the shock disappears and the new equilibrium price level is set.

In view of the theory, we find the shape of the estimated impulse response functions very convincing. The short- and long-run impacts, of course, differ across countries due to structural differences. A demand disturbance increases prices permanently, although the initial effect is much larger than the long run effect. Core shocks also increase output initially, but the effect dies out and the impulse response is close to zero, reflecting the imposed output-neutrality assumption.

The variance decomposition results are reported in Figures 11, 15, 19, 23, 27, 31, 35, 39 and 43. According to the definition of core inflation, its fluctuations are mainly explained by the core

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[20] By visual inspection, we find that the core inflation process is very much the same as the one reported by Blix (1995). Deviations of CPI inflation are substantial. Periods of under- and overestimation can be distinguished easily.

[21] We consider the GDP measure to be the more general proxy.

(demand) innovation for all countries. This observation is most accentuated for Italy (Figure 31) and the Netherlands (Figure 35). It is less pronounced for Austria (Figure 11), Germany (Figure 27), Belgium (Figure 15), France (Figure 23) and Sweden (Figure 39). Finland (Figure 19) and the United Kingdom (Figure 43) constitute exceptions, because core and non-core innovations explain more or less equal parts of CPI inflation forecast variance.

3.3 Trivariate SVAR

In a second step we differentiate monetary or LM shocks from real demand shocks. Both of these shocks were restricted not to have long-lasting effects on the level of output. This implies that both are core innovations, driving the underlying inflation process. The objective of the model extension is to investigate whether real aggregate demand and monetary innovations have similar effects on measured inflation. We also expect that the estimates for the inflation measures could be improved by the extension. We estimate a trivariate VAR system in GDP growth, Δy_t, the change in nominal interest rates, Δi_t, and in quarterly CPI inflation rates, Δp_t. The estimation results for all countries are summarised in Figures 1 to 9. The growth rates are calculated on a year-on-year basis. Again, the estimation period is 1970:1 to 1996:12. The values for 1997 and 1998 are forecasts. The system includes 3 lags, which is supported by various information criteria applied.[22] As previously indicated, this specification is consistent with y_t, i_t and p_t being I(1) (integrated of order one). Cointegration tests do not give evidence of cointegrating vectors.[23]

3.3.1 Core inflation versus CPI inflation

The estimation results for all countries are summarised again in Figures 1 to 9. Even though the Core CPI differentials differ somewhat from those obtained in the bivariate approach, the pattern of deviations closely matches the one of the previous results. In almost every case, the cyclical pattern of over- and underestimations is remarkably similar across both specifications.

For Austria, Belgium and Germany, the difference between the bivariate and the trivariate approach is negligible. For Finland, the Netherlands, Sweden and United Kingdom the deviations are minor. For France and Italy differences in the results are more important.

3.3.2 Impulse response functions and variance decompositions

The impulse response estimates for the trivariate VAR systems displayed in Figures 12, 16, 20, 24, 28, 32, 36, 40 and 44 reveal significant differences in the effects of real and monetary demand shocks on measured inflation. According to the theoretical background outlined above,[24] we expect the monetary policy or LM innovations to have negligible output and price effects for countries credibly pegging their exchange rate, and positive effects for countries with lesser credibility of the peg. Such "credibility effects" can only be found for Austria (Figure 12), Germany (Figure 28) and the Netherlands (Figure 36). As we observe negative price effects in the latter case, we might interpret this interest rate increase in the traditional manner as resulting from autonomous restrictive monetary measures. In all other countries monetary innovations increase output temporarily and prices even in the long-run.[25]

As in the bivariate case, we estimated variance decompositions for each country. The results are shown in Figure 13 (Austria), 17 (Belgium), 21 (Finland), 25 (France), 29 (Germany),

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22 See footnote 17.

23 See footnote 16.

24 See also footnote 11.

25 Due to our identifying restrictions, we do not allow for long-run output effects of a nominal interest rate shock.

33 (Italy), 37 (Netherlands), 41 (Sweden) and 45 (United Kingdom). We can not fully confirm the findings by Dewachter and Lustig (1997). We have already touched upon the problem of differences in results when describing the impulse response functions for the trivariate case: Interpreting their variance decompositions, Dewachter and Lustig (1997) discovered that the inflationary process is mainly driven by monetary shocks, rather than real (core) shocks. In the long run, 75% to 95% of the variability in measured inflation are accounted for by monetary innovations. Referring to the respective figures, they conclude that inflation is really a monetary phenomenon. According to our estimates, we can share their opinion on inflation being essentially demand driven, but we cannot support the judgement of inflation being a purely monetary phenomenon.

4. Does monetary policy co-ordination enhance inflation convergence? A correlation analysis

In Section 3 we calculated indicators for the underlying inflation process. These core inflation indicators are considered to be more relevant assessing the sustainability of a country's inflation performance than the conventional CPI inflation measure. For the assessment of the ECB's single monetary policy, it is important to know whether there are common trends or common cycles in inflation performance of EU member states. We will address this issue stepwise.

First, we start by a cross correlation analysis involving the CPI inflation and core inflation indicators of the selected countries, whereby we seek to answer the following questions:

Hypothesis 1: We expect that the correlation coefficient between inflation indicators is higher if the country belongs to the "core group" (Austria, Belgium, France, Germany and the Netherlands) rather than to the "periphery group" (Finland, Italy, Sweden and the United Kingdom).[26]

Hypothesis 2: We also expect that the correlation coefficient between the CPI inflation measures is lower than that between the core inflation measures.

Hypothesis 3: We expect that the correlation coefficients are higher in the 1990s due to enhanced monetary policy co-ordination and economic integration than the figures calculated for the 1970s and 1980s, respectively.

As far as the first hypothesis is concerned, we find weak evidence that the inflation performance correlation among ERM countries is closer than among the "periphery group" (see Table 1 in Appendix B). The results are distorted due to the rather "ad hoc" definition of the groups (e.g. relatively high correlation coefficients of France/Italy and rather low coefficients of Austria/France).

Similar results were obtained for the second hypothesis (see Table 1). In most of the cases it seems that especially the correlation coefficient for the core inflation indicators calculated by the bivariate decomposition, "Core Infl. (2)", is slightly higher than between actual inflation rates; the differences are not significant, though.

As to the third hypothesis, the results do not allow us to give a clear answer (cf. Table 1).

Further analysis is required. Cointegration analysis of core inflation series could cast some more light on the existence of common inflation trends in the EU.

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[26] We define "core countries" as the ones that have (at least during most of the estimation period) been tying their currency explicitly to the Deutsche mark, and Germany itself.

Conclusions

We calculated core inflation indicators for Austria, Belgium, Finland, France, Germany, Italy, the Netherlands, Sweden and the United Kingdom in a structural VAR framework applying long-run identification schemes similar to the ones proposed by Quah and Vahey (1995). As also suggested by their work we included a third variable in the VAR system, short-term nominal interest rates, which we assumed to capture the effects of monetary disturbances in the system. Contrasting the results (when applicable) to those of Blix (1995) and Dewachter and Lustig (1997), they differ in many respects for obvious reasons: First of all, we used quarterly instead of monthly data, because we included GDP instead of industrial production data in our analysis. Secondly, especially in the trivariate case, we used a different identification scheme (e.g., both Blix (1995) and Dewachter and Lustig (1997) included cointegrating restrictions motivated by economic theory). Specifically, we use changes of prices instead of changes in inflation in our estimations and impose respective long-term restrictions in this context. The analysis bears on an IS–LM/AS–AD framework for small open economies and/or countries with fixed exchange rate regimes.

Dewachter and Lustig (1997) find that the inflation process is mainly driven by monetary shocks, rather than demand shocks. Hence, they conclude that inflation is a monetary phenomenon. According to our estimates, we find that inflation is essentially demand-driven, but our results at this stage do not support their view that inflation is a purely monetary phenomenon.

A cross correlation analysis completes the paper, this exercise being a first attempt to address the question about the existence of common inflation trends in EU countries. Future research should aim for an in-depth analysis of common trends and cycles among EU inflation measures.

Appendix A: Confidence bands of impulse response functions

In order to report two-standard error bands in the graphs of the impulse response functions as shown below we apply a Monte-Carlo approach. Although there is a common procedure for the "traditional" VARs that use short-term restrictions to identify the structural shocks, the calculation of the error bands for VARs using long-run restrictions are, as of now, not common knowledge among model builders. So far, also an analytical approach – which is given by Lütkepohl (1993, p. 313ff) for "traditional" VARs – has not been finally designed in the context of long-run identifying restrictions.[27] Here we use a slightly modified version of a technique expounded in, e.g., Mélitz and Weber (1996).[28]

If we write the VAR as:

$$y_t = (\mathrm{I} \otimes x_t)\beta + u_t$$

where $\otimes$ is the Kronecker product, x_t is the vector of lagged y_{it}'s ($i = 1,2,...m$), β is a vector containing the stacked version of the structural VAR lag polynomial matrices, $A(L)$, and u_t is i.i.d. with distribution $\mathrm{N}(0,\Sigma)$. The OLS estimates of β and Σ are denoted by b and Z. Assuming that the prior distribution of β is $f(\beta,\Sigma) \propto |\Sigma|^{-(n+1)/2}$, the posterior distribution of β, conditional on Σ, is $\mathrm{N}(b, \Sigma \otimes (x'x)^{-1})$ and the distribution of Σ^{-1} is $\mathrm{Wishart}((TZ)^{-1}, T)$ with T as sample size.

First and second moments for the impulse responses (the moving average representation) can be computed by drawing q times[29] from the above distribution for β and Σ, inverting the VAR, calculating each time[30] the innovation-orthogonalising matrix E_0^{-1} (as shown in the text) and conditional on that calculating the mean and the variance impulse responses (moving average parameters).

In order to derive standard errors for the accumulated impulse responses as shown in the graphs (for "level series"), we accumulate the impulses of each of the q draws for every impulse step period p, calculate their variance over the q draws and then adjust this variance in each impulse step, multiplying it by p^{-1}. The standard errors are then given by the square root of the resulting adjusted variances. We perform this adjustment referring to the fact that the identifying restrictions are imposed on the long-run moving average parameters, i.e. the accumulations of the moving average parameters derived from the estimated model with differenced series, and any variance of the accumulated parameters at step p has to be treated as sample variance of the parameters up to step p.

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27 But see the suggestion by Vlaar (1997).

28 For the calculations we modify a RATS program procedure given in Doan (1992, p.10-5).

29 We used $q = 300$ for our calculations.

30 Here we differ from the approach as given in Melitz and Weber (1996); they perform the calculations conditional on E_0^{-1} as derived from the initial estimation.

Appendix B: Tables and graphs

Table 1

Cross correlations of inflation series between countries

		Belgium	Finland	France	Germany	Italy	Netherlands	Sweden	U. Kingdom
Austria	Actual Infl.	0.88	0.83	0.76	0.82	0.72	0.94	0.57	0.73
1971:1-96:4	Core Infl. (2)	0.92	0.82	0.72	0.84	0.73	0.96	0.61	0.68
	Core Infl. (3)	0.81	0.74	0.76	0.82	0.85	0.92	0.71	0.36
	Actual-Core (2)	0.56	0.32	0.59	0.42	0.12	0.57	0.04	0.50
	Actual-Core (3)	0.14	0.11	0.58	0.40	0.07	0.36	0.34	0.29
1971:1-80:4	Actual Infl.	0.85	0.73	0.35	0.73	0.33	0.92	0.10	0.44
	Core Infl. (2)	0.85	0.65	0.14	0.81	0.36	0.92	-0.08	0.32
	Core Infl. (3)	0.59	0.66	0.87	-0.12	0.94	0.93	0.51	-0.24
	Actual-Core (2)	0.64	0.33	0.74	0.60	0.35	0.69	-0.07	0.67
	Actual-Core (3)	-0.51	0.95	0.67	-0.01	-0.88	0.84	0.97	0.39
1981:1-90:4	Actual Infl.	0.82	0.81	0.86	0.84	0.83	0.93	0.74	0.62
	Core Infl. (2)	0.86	0.88	0.84	0.90	0.86	0.96	0.92	0.07
	Core Infl. (3)	0.85	0.85	0.81	0.89	0.86	0.95	0.83	0.00
	Actual-Core (2)	0.49	0.27	0.70	0.35	-0.14	0.49	-0.25	0.37
	Actual-Core (3)	0.24	0.23	0.76	0.34	-0.19	0.34	-0.04	0.20
1990:1-96:4	Actual Infl.	0.65	0.57	0.37	0.80	0.36	0.65	0.42	0.22
	Core Infl. (2)	0.69	0.06	0.61	0.86	0.39	0.73	0.54	-0.02
	Core Infl. (3)	0.64	0.56	0.64	0.84	0.24	0.69	0.44	0.11
	Actual-Core (2)	0.45	0.43	0.45	0.46	0.21	0.71	0.53	0.23
	Actual-Core (3)	0.07	-0.16	0.46	0.52	0.22	0.31	0.69	0.29
Belgium	Actual Infl.	1.00	0.89	0.84	0.68	0.82	0.85	0.63	0.75
1971:1-96:4	Core Infl. (2)	1.00	0.90	0.82	0.72	0.79	0.89	0.67	0.79
	Core Infl. (3)	1.00	0.84	0.88	0.75	0.83	0.76	0.74	0.43
	Actual-Core (2)	1.00	0.10	0.67	0.30	-0.11	0.34	0.01	0.31
	Actual-Core (3)	1.00	-0.68	0.41	0.08	-0.25	-0.36	-0.31	-0.09
1971:1-80:4	Actual Infl.	1.00	0.89	0.56	0.48	0.51	0.82	0.24	0.64
	Core Infl. (2)	1.00	0.90	0.38	0.52	0.57	0.85	0.09	0.56
	Core Infl. (3)	1.00	0.89	0.48	0.27	0.82	0.74	0.67	-0.61
	Actual-Core (2)	1.00	0.50	0.84	0.60	0.39	0.51	0.07	0.80
	Actual-Core (3)	1.00	-0.61	-0.96	0.39	0.61	-0.49	-0.37	0.43
1981:1-90:4	Actual Infl.	1.00	0.81	0.90	0.83	0.92	0.85	0.66	0.41
	Core Infl. (2)	1.00	0.92	0.83	0.89	0.83	0.84	0.80	-0.08
	Core Infl. (3)	1.00	0.89	0.89	0.90	0.88	0.87	0.78	0.02
	Actual-Core (2)	1.00	-0.07	0.71	0.17	-0.28	0.42	-0.18	0.38
	Actual-Core (3)	1.00	-0.71	0.43	-0.02	-0.40	-0.28	-0.35	0.12
1990:1-96:4	Actual Infl.	1.00	0.81	0.75	0.38	0.42	0.56	0.75	0.61
	Core Infl. (2)	1.00	0.26	0.94	0.73	0.17	0.55	0.70	0.15
	Core Infl. (3)	1.00	0.47	0.90	0.78	-0.18	0.38	0.57	0.17
	Actual-Core (2)	1.00	-0.38	0.90	0.49	-0.44	0.22	0.04	-0.62
	Actual-Core (3)	1.00	-0.85	0.79	0.31	-0.61	-0.57	-0.34	-0.83
Finland	Actual Infl.		1.00	0.86	0.63	0.84	0.82	0.77	0.88
1971:1-96:4	Core Infl. (2)		1.00	0.91	0.58	0.83	0.80	0.77	0.89
	Core Infl. (3)		1.00	0.93	0.49	0.85	0.66	0.80	0.70
	Actual-Core (2)		1.00	0.23	0.31	0.05	0.31	0.44	0.60
	Actual-Core (3)		1.00	-0.22	-0.08	0.44	0.44	0.30	0.38
1971:1-80:4	Actual Infl.		1.00	0.58	0.40	0.58	0.73	0.43	0.77
	Core Infl. (2)		1.00	0.40	0.27	0.60	0.64	0.12	0.69
	Core Infl. (3)		1.00	0.36	-0.19	0.79	0.73	0.58	-0.85
	Actual-Core (2)		1.00	0.61	0.47	-0.31	0.19	0.29	0.66
	Actual-Core (3)		1.00	0.79	0.00	-0.82	0.70	0.95	0.33
1981:1-90:4	Actual Infl.		1.00	0.90	0.91	0.87	0.88	0.78	0.68
	Core Infl. (2)		1.00	0.95	0.89	0.95	0.92	0.83	0.09
	Core Infl. (3)		1.00	0.94	0.84	0.91	0.92	0.73	0.18
	Actual-Core (2)		1.00	0.38	0.62	0.61	0.50	0.55	0.21
	Actual-Core (3)		1.00	0.06	0.19	0.36	0.44	0.55	-0.01
1990:1-96:4	Actual Infl.		1.00	0.88	0.33	0.74	0.40	0.90	0.84
	Core Infl. (2)		1.00	0.44	-0.22	0.35	-0.13	0.68	0.96
	Core Infl. (3)		1.00	0.50	0.36	0.01	0.16	0.62	0.50
	Actual-Core (2)		1.00	-0.38	-0.09	0.49	0.50	0.58	0.86
	Actual-Core (3)		1.00	-0.71	-0.39	0.53	0.60	0.16	0.66

Table 2

Cross correlations of inflation series between countries (continued)

		Belgium	Finland	France	Germany	Italy	Netherlands	Sweden	U. Kingdom
France	Actual Infl.			1.00	0.63	0.95	0.75	0.75	0.77
1971:1-96:4	Core Infl. (2)			1.00	0.57	0.88	0.69	0.86	0.77
	Core Infl. (3)			1.00	0.56	0.91	0.70	0.84	0.73
	Actual-Core (2)			1.00	0.47	-0.16	0.48	-0.04	0.45
	Actual-Core (3)			1.00	0.44	-0.38	0.42	0.05	0.24
1971:1-80:4	Actual Infl.			1.00	0.14	0.89	0.15	0.58	0.66
	Core Infl. (2)			1.00	-0.25	0.88	-0.02	0.82	0.37
	Core Infl. (3)			1.00	0.20	0.85	0.89	0.61	0.17
	Actual-Core (2)			1.00	0.67	0.38	0.58	0.03	0.81
	Actual-Core (3)			1.00	-0.34	-0.73	0.53	0.57	-0.26
1981:1-90:4	Actual Infl.			1.00	0.86	0.96	0.92	0.65	0.52
	Core Infl. (2)			1.00	0.83	0.98	0.90	0.73	0.12
	Core Infl. (3)			1.00	0.83	0.96	0.90	0.69	0.19
	Actual-Core (2)			1.00	0.35	-0.10	0.60	-0.07	0.57
	Actual-Core (3)			1.00	0.33	-0.18	0.46	0.16	0.34
1990:1-96:4	Actual Infl.			1.00	0.24	0.82	0.31	0.86	0.84
	Core Infl. (2)			1.00	0.66	0.35	0.37	0.81	0.32
	Core Infl. (3)			1.00	0.74	0.15	0.31	0.81	0.48
	Actual-Core (2)			1.00	0.71	-0.54	0.27	-0.19	-0.67
	Actual-Core (3)			1.00	0.68	-0.58	-0.28	-0.18	-0.62
Germany	Actual Infl.				1.00	0.56	0.84	0.44	0.57
1971:1-96:4	Core Infl. (2)				1.00	0.56	0.85	0.52	0.37
	Core Infl. (3)				1.00	0.61	0.84	0.53	0.09
	Actual-Core (2)				1.00	0.18	0.51	-0.01	0.34
	Actual-Core (3)				1.00	0.04	0.21	0.32	0.22
1971:1-80:4	Actual Infl.				1.00	0.02	0.67	-0.12	0.23
	Core Infl. (2)				1.00	-0.10	0.80	-0.34	-0.03
	Core Infl. (3)				1.00	0.08	-0.01	0.02	0.44
	Actual-Core (2)				1.00	0.38	0.49	-0.16	0.77
	Actual-Core (3)				1.00	0.48	0.06	0.19	0.82
1981:1-90:4	Actual Infl.				1.00	0.85	0.93	0.78	0.79
	Core Infl. (2)				1.00	0.82	0.92	0.89	0.10
	Core Infl. (3)				1.00	0.79	0.92	0.85	0.15
	Actual-Core (2)				1.00	0.42	0.62	0.18	0.11
	Actual-Core (3)				1.00	0.30	0.26	0.34	0.00
1990:1-96:4	Actual Infl.				1.00	0.21	0.74	0.16	-0.07
	Core Infl. (2)				1.00	0.27	0.69	0.45	-0.29
	Core Infl. (3)				1.00	0.11	0.69	0.48	0.02
	Actual-Core (2)				1.00	-0.47	0.42	-0.12	-0.34
	Actual-Core (3)				1.00	-0.44	-0.04	0.18	-0.08
Italy	Actual Infl.					1.00	0.70	0.76	0.77
1971:1-96:4	Core Infl. (2)					1.00	0.72	0.80	0.70
	Core Infl. (3)					1.00	0.84	0.76	0.57
	Actual-Core (2)					1.00	0.24	0.17	0.21
	Actual-Core (3)					1.00	0.07	0.43	0.27
1971:1-80:4	Actual Infl.					1.00	0.15	0.64	0.66
	Core Infl. (2)					1.00	0.25	0.74	0.66
	Core Infl. (3)					1.00	0.97	0.66	-0.36
	Actual-Core (2)					1.00	0.31	-0.30	0.24
	Actual-Core (3)					1.00	-0.73	-0.76	0.05
1981:1-90:4	Actual Infl.					1.00	0.90	0.67	0.47
	Core Infl. (2)					1.00	0.90	0.78	0.05
	Core Infl. (3)					1.00	0.91	0.73	0.09
	Actual-Core (2)					1.00	0.22	0.77	0.03
	Actual-Core (3)					1.00	0.33	0.48	0.02
1990:1-96:4	Actual Infl.					1.00	0.20	0.82	0.78
	Core Infl. (2)					1.00	0.15	0.64	0.34
	Core Infl. (3)					1.00	-0.06	0.31	0.26
	Actual-Core (2)					1.00	0.19	0.65	0.71
	Actual-Core (3)					1.00	0.46	0.58	0.66

Table 3

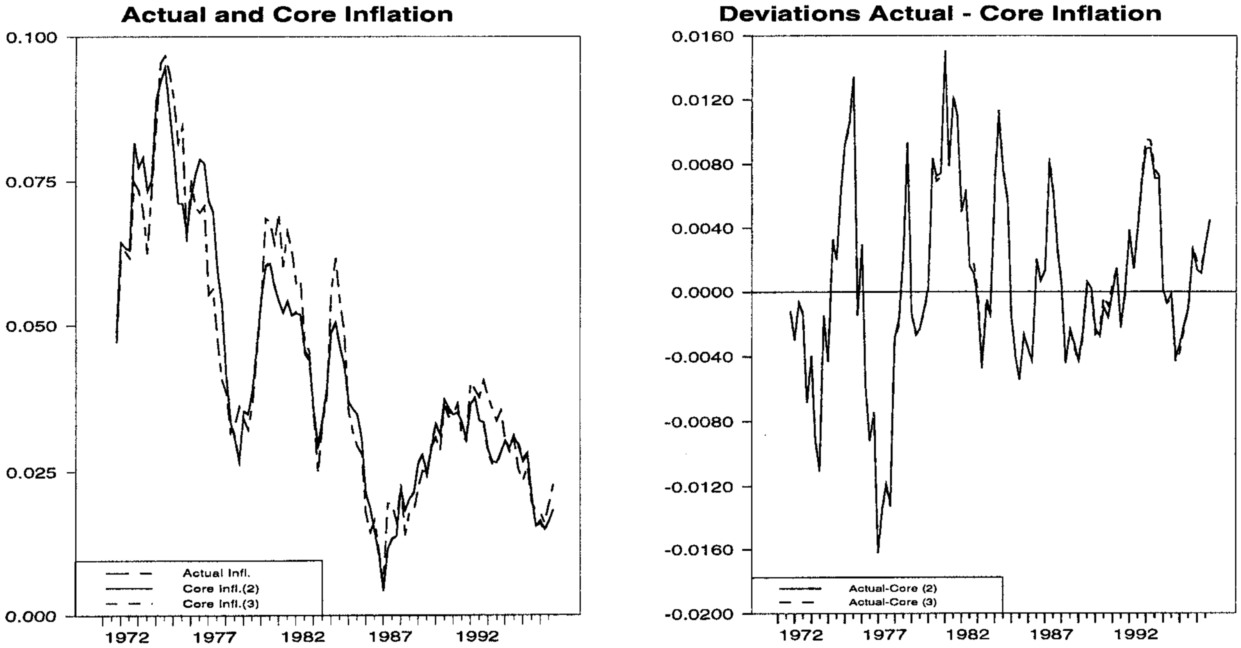
Cross correlations of inflation series between countries (continued)

		Belgium	Finland	France	Germany	Italy	Netherlands	Sweden	U. Kingdom
Netherlands	Actual Infl.						1.00	0.54	0.76
1971:1-96:4	Core Infl. (2)						1.00	0.57	0.67
	Core Infl. (3)						1.00	0.58	0.28
	Actual-Core (2)						1.00	0.23	0.46
	Actual-Core (3)						1.00	0.36	0.34
1971:1-80:4	Actual Infl.						1.00	0.02	0.44
	Core Infl. (2)						1.00	-0.18	0.32
	Core Infl. (3)						1.00	0.77	-0.28
	Actual-Core (2)						1.00	0.12	0.55
	Actual-Core (3)						1.00	0.78	0.29
1981:1-90:4	Actual Infl.						1.00	0.78	0.73
	Core Infl. (2)						1.00	0.88	0.18
	Core Infl. (3)						1.00	0.80	0.17
	Actual-Core (2)						1.00	0.22	0.28
	Actual-Core (3)						1.00	0.34	0.12
1990:1-96:4	Actual Infl.						1.00	0.26	0.12
	Core Infl. (2)						1.00	0.37	-0.08
	Core Infl. (3)						1.00	0.11	-0.11
	Actual-Core (2)						1.00	0.44	0.36
	Actual-Core (3)						1.00	0.43	0.61
Sweden	Actual Infl.							1.00	0.71
1971:1-96:4	Core Infl. (2)							1.00	0.66
	Core Infl. (3)							1.00	0.75
	Actual-Core (2)							1.00	0.18
	Actual-Core (3)							1.00	0.42
1971:1-80:4	Actual Infl.							1.00	0.53
	Core Infl. (2)							1.00	0.27
	Core Infl. (3)							1.00	-0.27
	Actual-Core (2)							1.00	0.07
	Actual-Core (3)							1.00	0.57
1981:1-90:4	Actual Infl.							1.00	0.72
	Core Infl. (2)							1.00	0.03
	Core Infl. (3)							1.00	0.07
	Actual-Core (2)							1.00	0.25
	Actual-Core (3)							1.00	0.29
1990:1-96:4	Actual Infl.							1.00	0.83
	Core Infl. (2)							1.00	0.65
	Core Infl. (3)							1.00	0.80
	Actual-Core (2)							1.00	0.56
	Actual-Core (3)							1.00	0.71

Note: The numbers in parentheses indicate the results of the bivariate (output, inflation) and trivariate (output, interest rate and inflation) model, respectively.

Figure 1

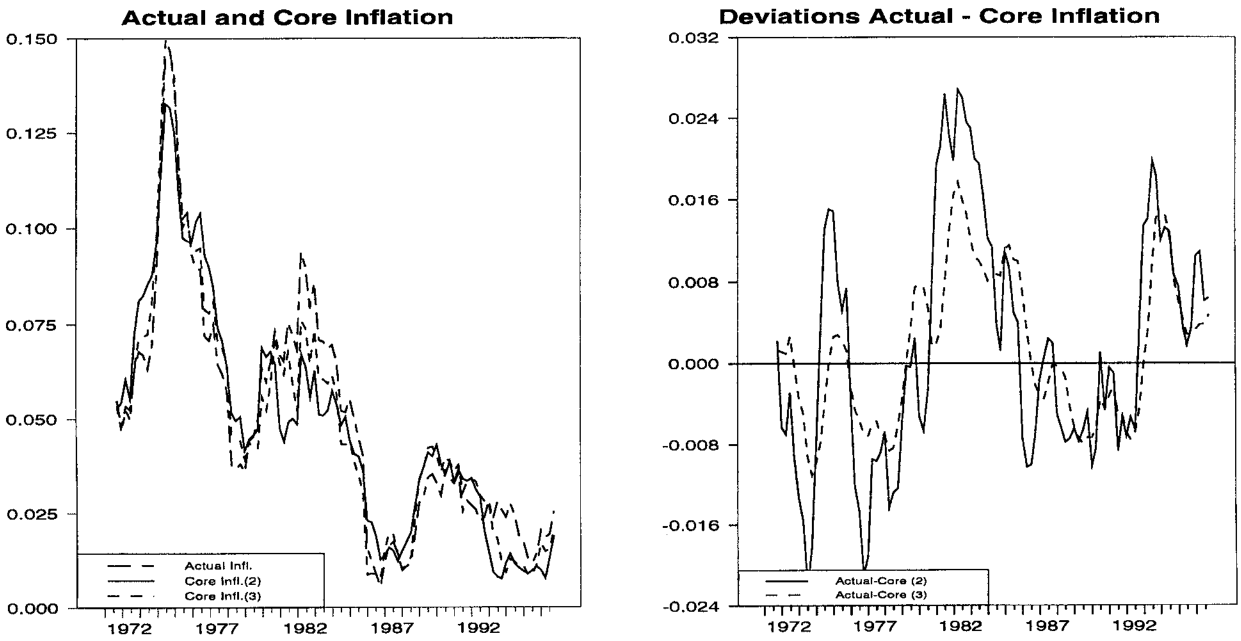
Inflation and core inflation in Austria



Note: The numbers in parentheses indicate the results of the bivariate (output, inflation) and trivariate (output, interest rate and inflation) model, respectively.

Figure 2

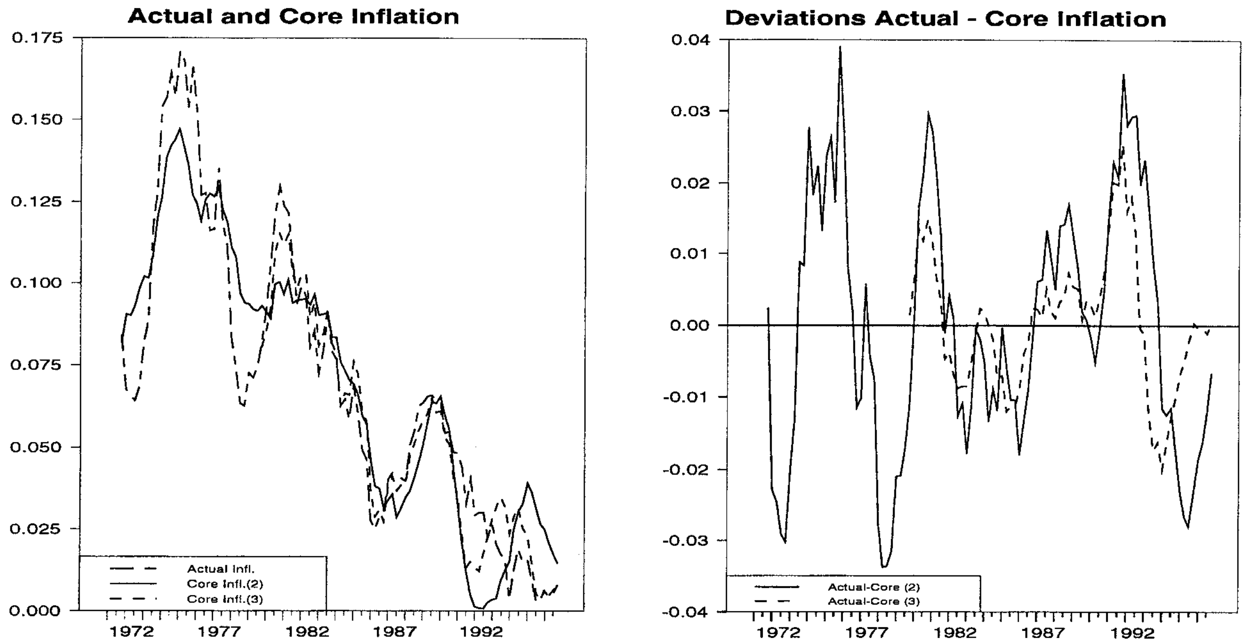
Inflation and core inflation in Belgium



Note: The numbers in parentheses indicate the results of the bivariate (output, inflation) and trivariate (output, interest rate and inflation) model, respectively.

Figure 3

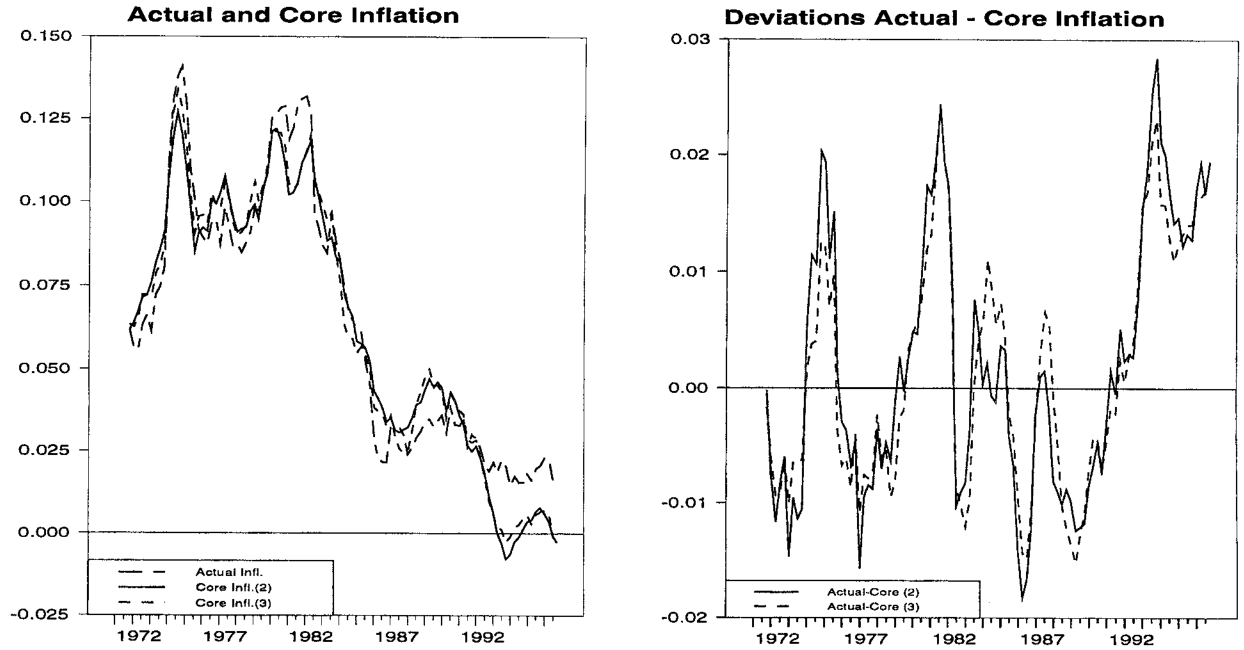
Inflation and core inflation in Finland



Note: The numbers in parentheses indicate the results of the bivariate (output, inflation) and trivariate (output, interest rate and inflation) model, respectively.

Figure 4

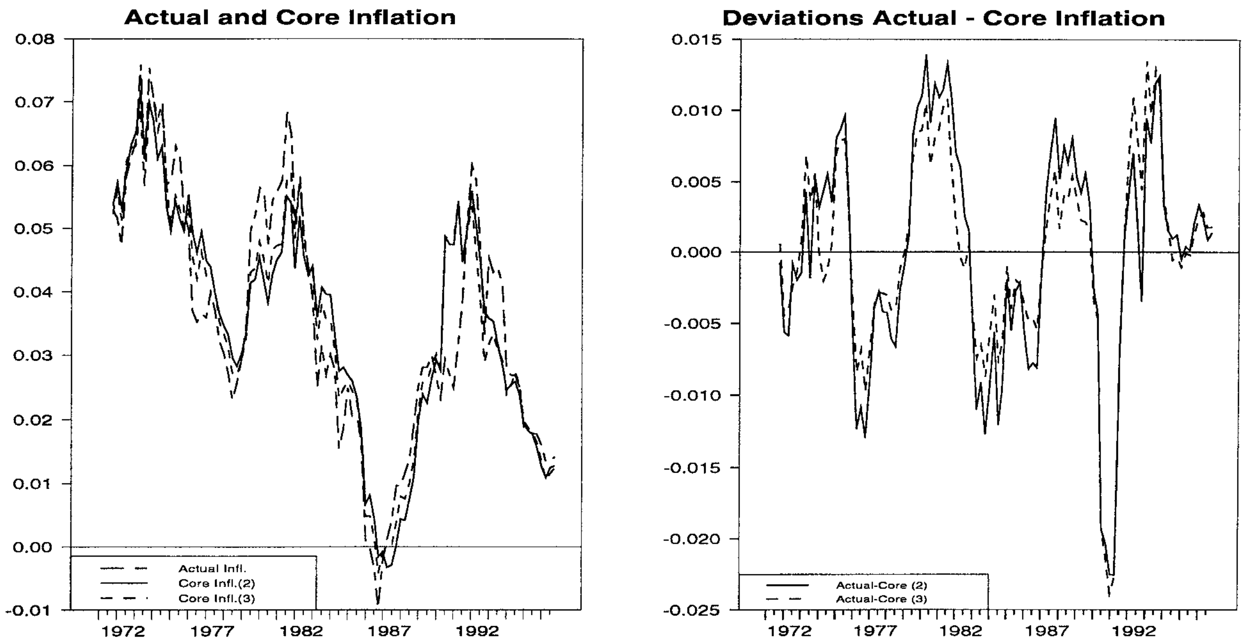
Inflation and core inflation in France



Note: The numbers in parentheses indicate the results of the bivariate (output, inflation) and trivariate (output, interest rate and inflation) model, respectively.

Figure 5

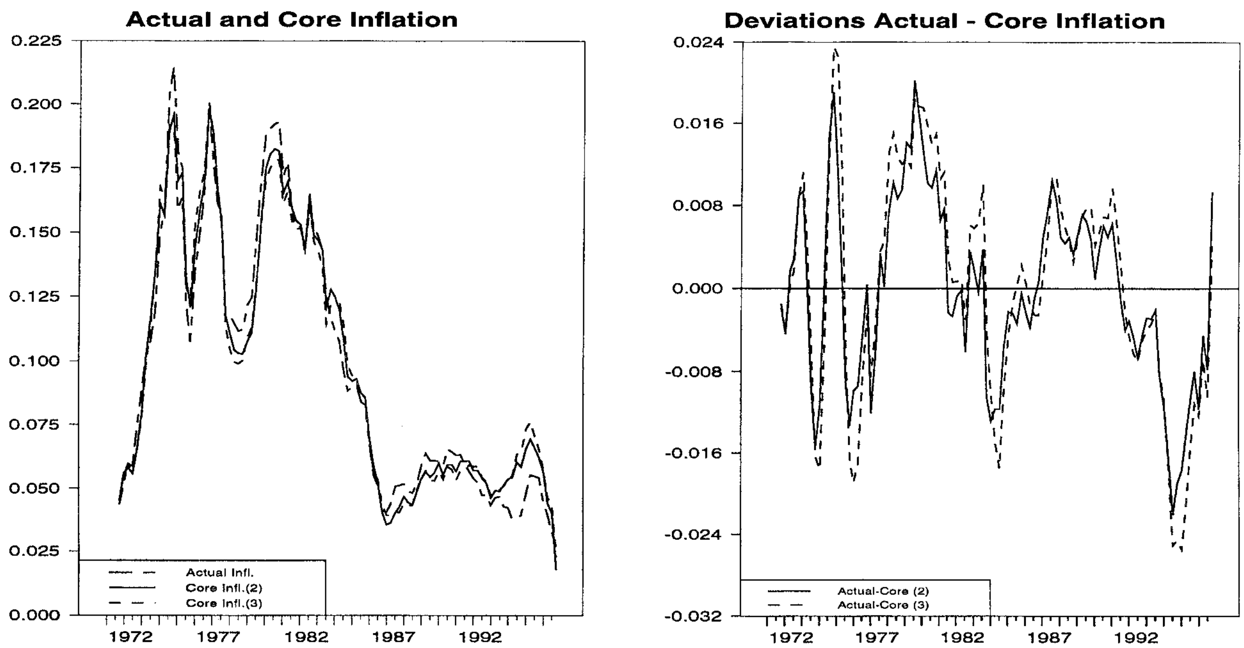
Inflation and core inflation in Germany



Note: The numbers in parentheses indicate the results of the bivariate (output, inflation) and trivariate (output, interest rate and inflation) model, respectively.

Figure 6

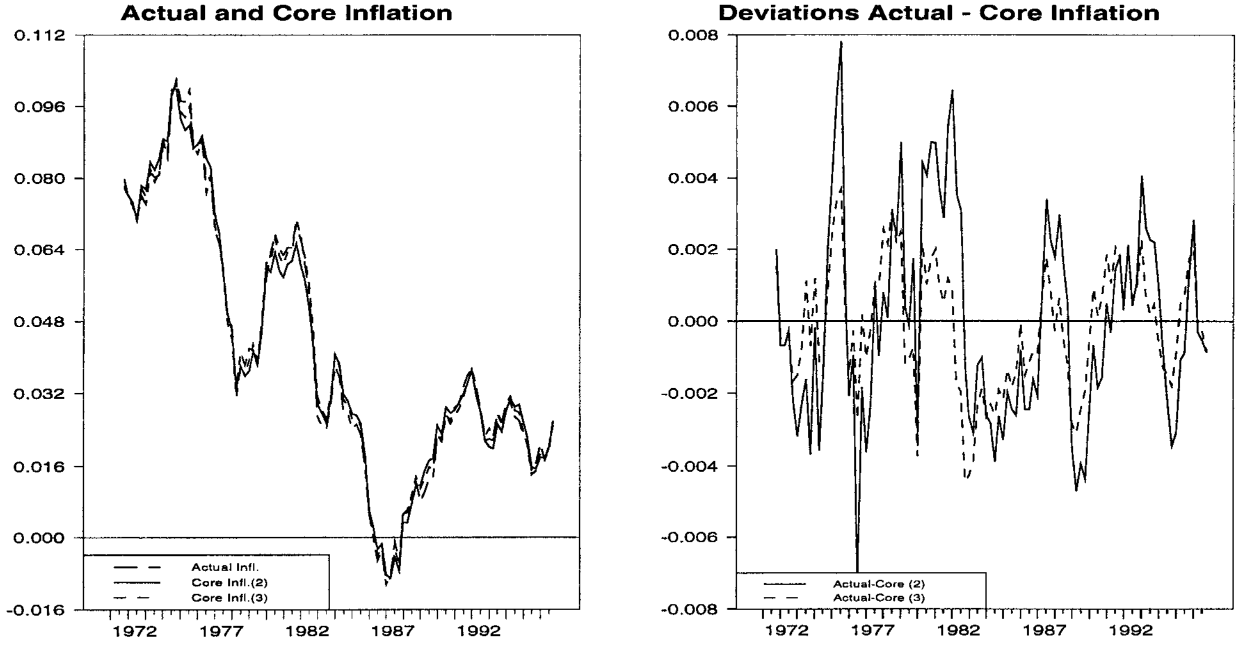
Inflation and core inflation in Italy



Note: The numbers in parentheses indicate the results of the bivariate (output, inflation) and trivariate (output, interest rate and inflation) model, respectively.

Figure 7

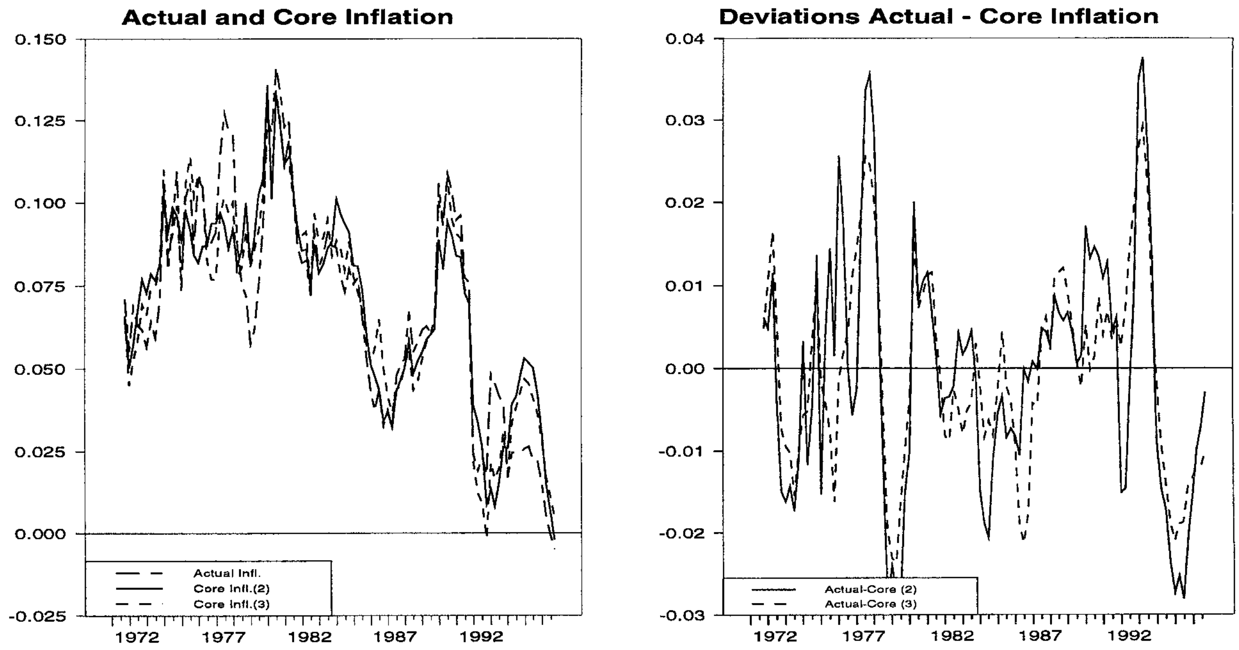
Inflation and core inflation in the Netherlands



Note: The numbers in parentheses indicate the results of the bivariate (output, inflation) and trivariate (output, interest rate and inflation) model, respectively.

Figure 8

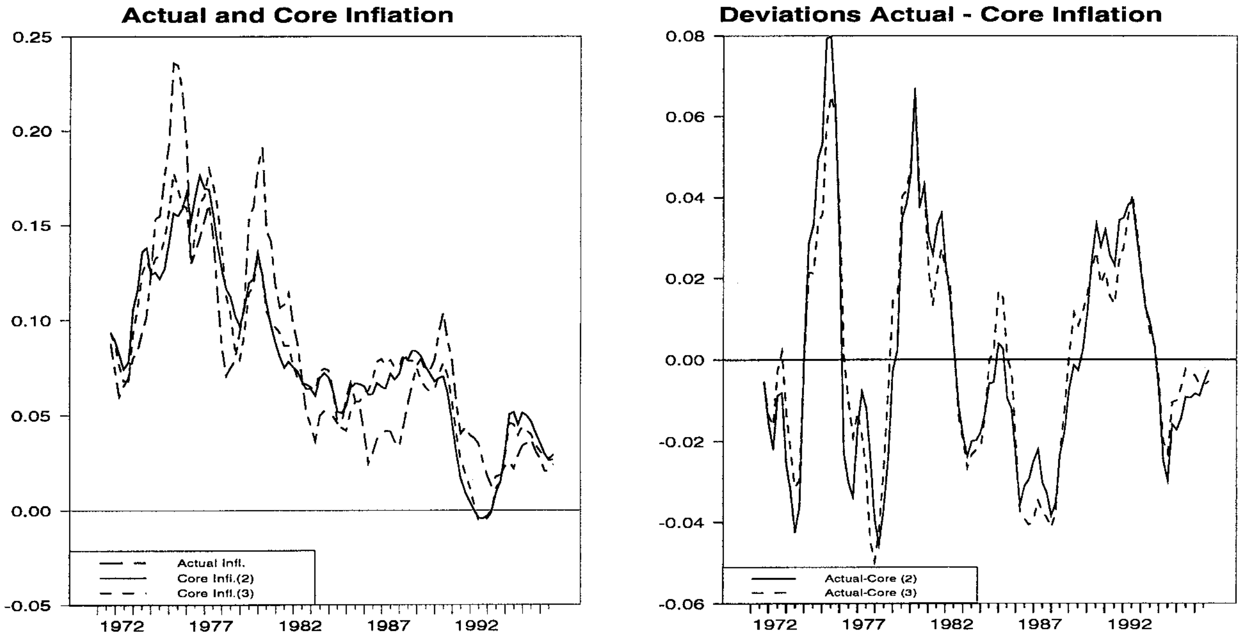
Inflation and core inflation in Sweden



Note: The numbers in parentheses indicate the results of the bivariate (output, inflation) and trivariate (output, interest rate and inflation) model, respectively.

Figure 9

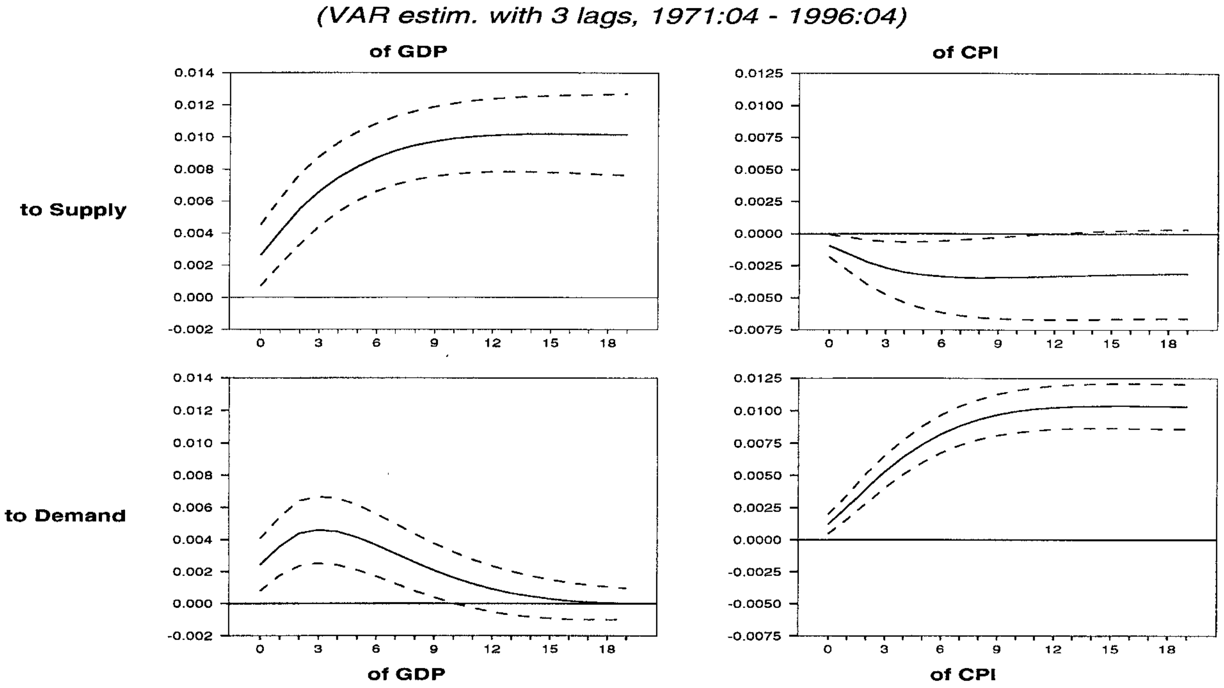
Inflation and core inflation in the United Kingdom



Note: The numbers in parentheses indicate the results of the bivariate (output, inflation) and trivariate (output, interest rate and inflation) model, respectively.

Figure 10

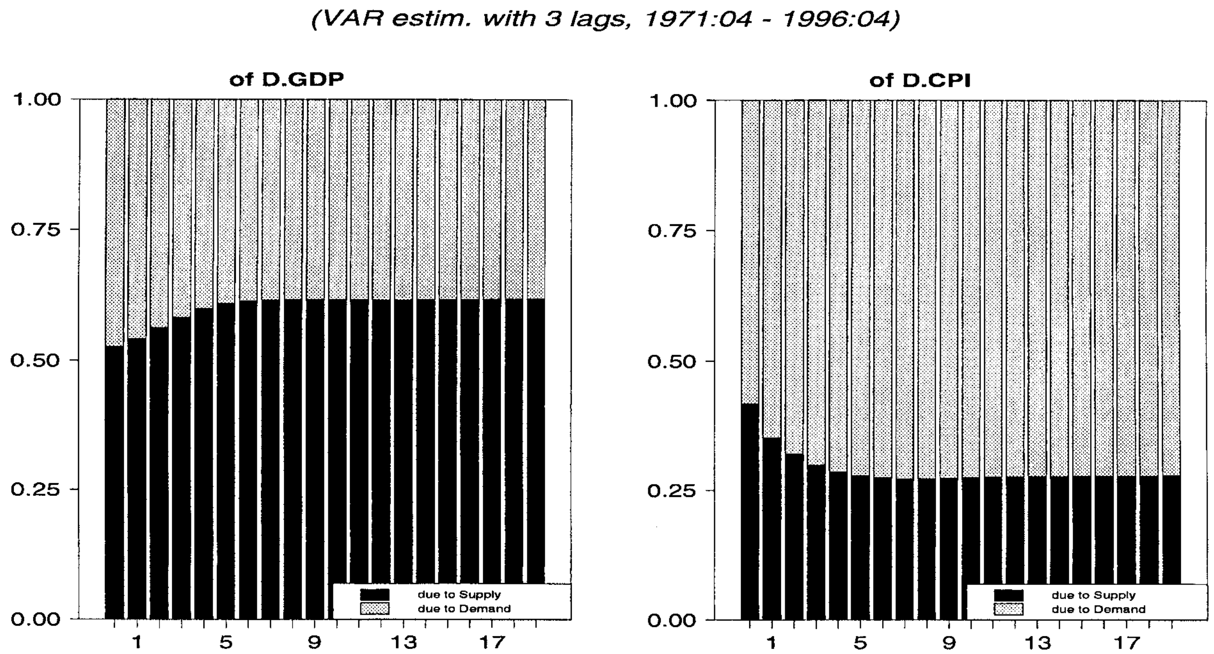
Impulse response functions (bivariate model) – Austria



Note: Results are those of the bivariate model (output and inflation). The solid lines indicate the estimated and accumulated response to the respective first period structural unit shock, dashed lines above and below are the upper and lower two standard deviation bounds computed from a simulation as described in Appendix A.

Figure 11

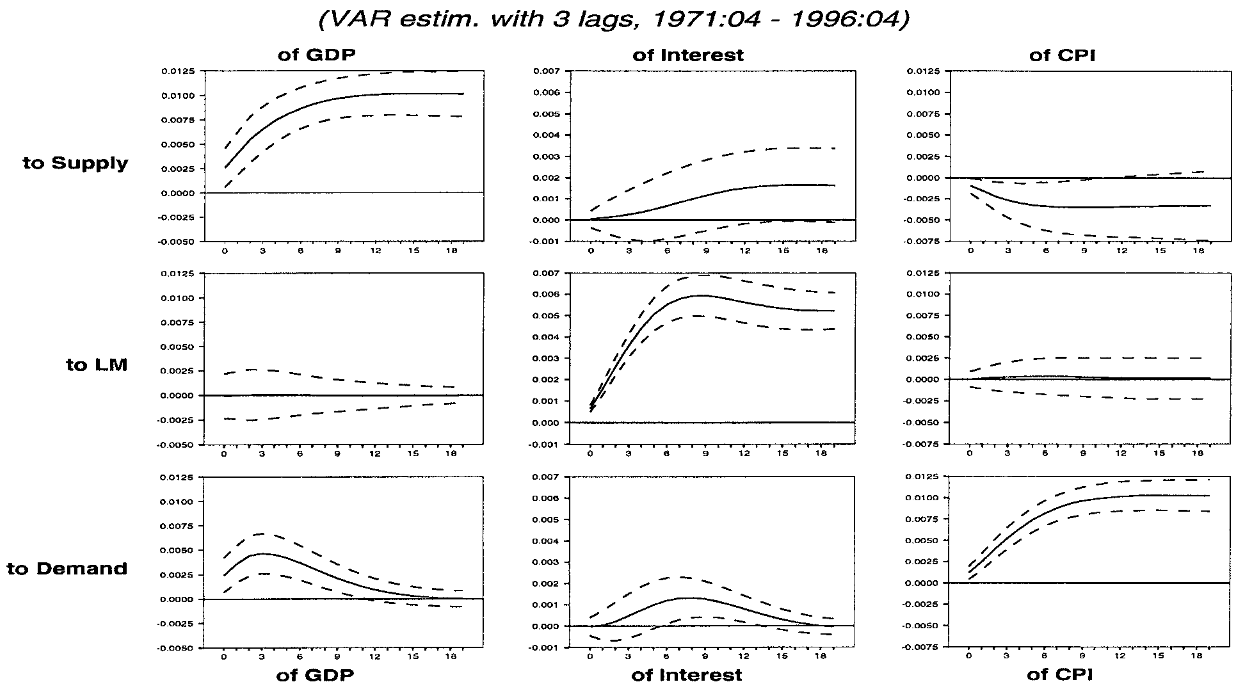
Variance decompositions (bivariate model) – Austria



Note: Results are those of the bivariate model (output and inflation). The heights of the respective bars indicate the relative contribution of a specific structural shock to the forecast error variance of the respective series.

Figure 12

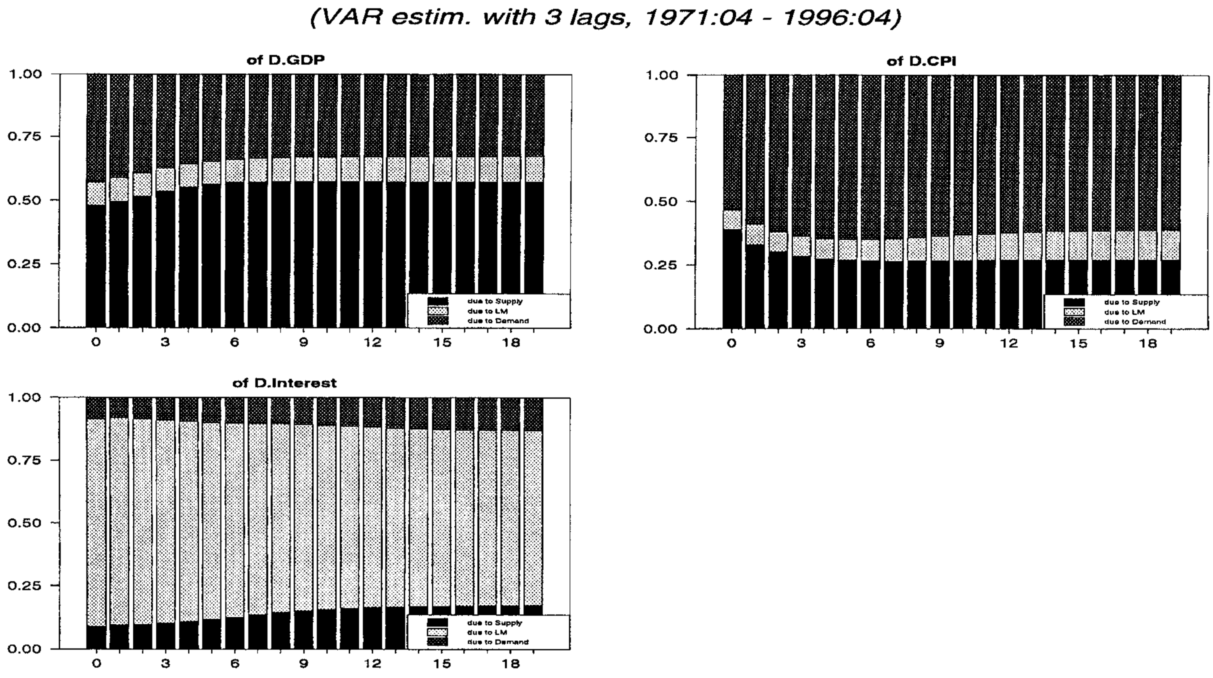
Impulse response functions (trivariate model) – Austria



Note: Results are those of the trivariate model (output, interest rate and inflation). The solid lines indicate the estimated and accumulated response to the respective first period structural unit shock, dashed lines above and below are the upper and lower two standard deviation bounds computed from a simulation as described in Appendix A.

Figure 13

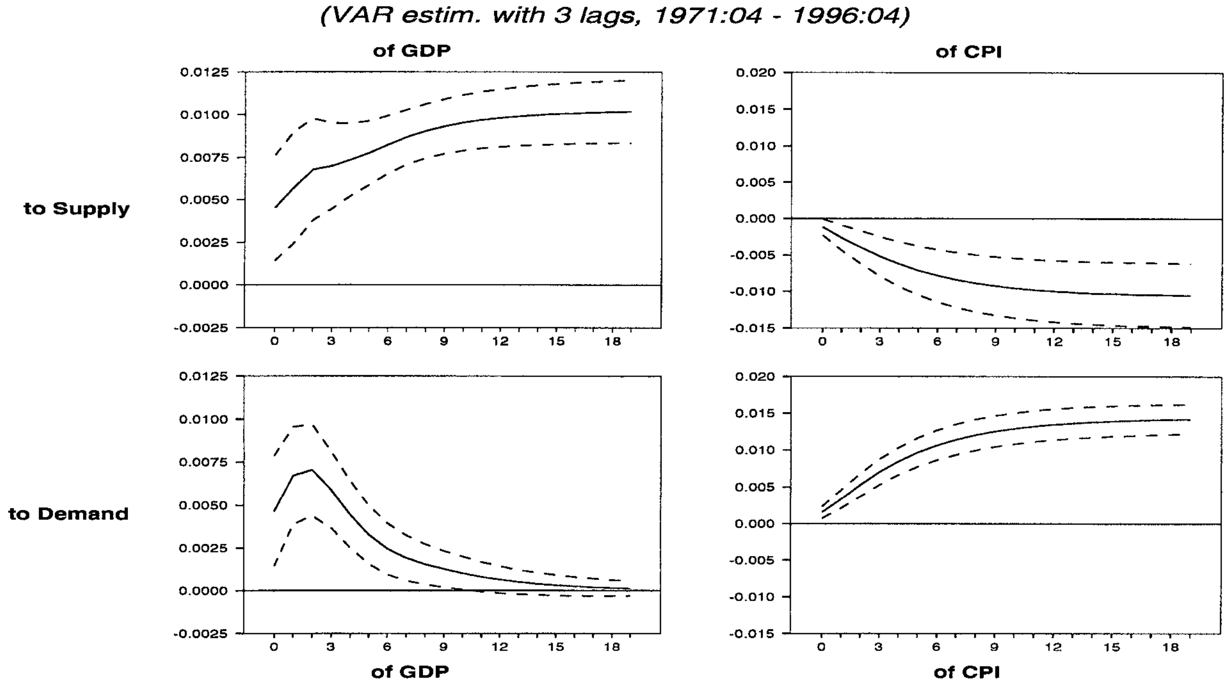
Variance decompositions (trivariate model) – Austria



Note: Results are those of the trivariate model (output, interest rate and inflation). The heights of the respective bars indicate the relative contribution of a specific structural shock to the forecast error variance of the respective series.

Figure 14

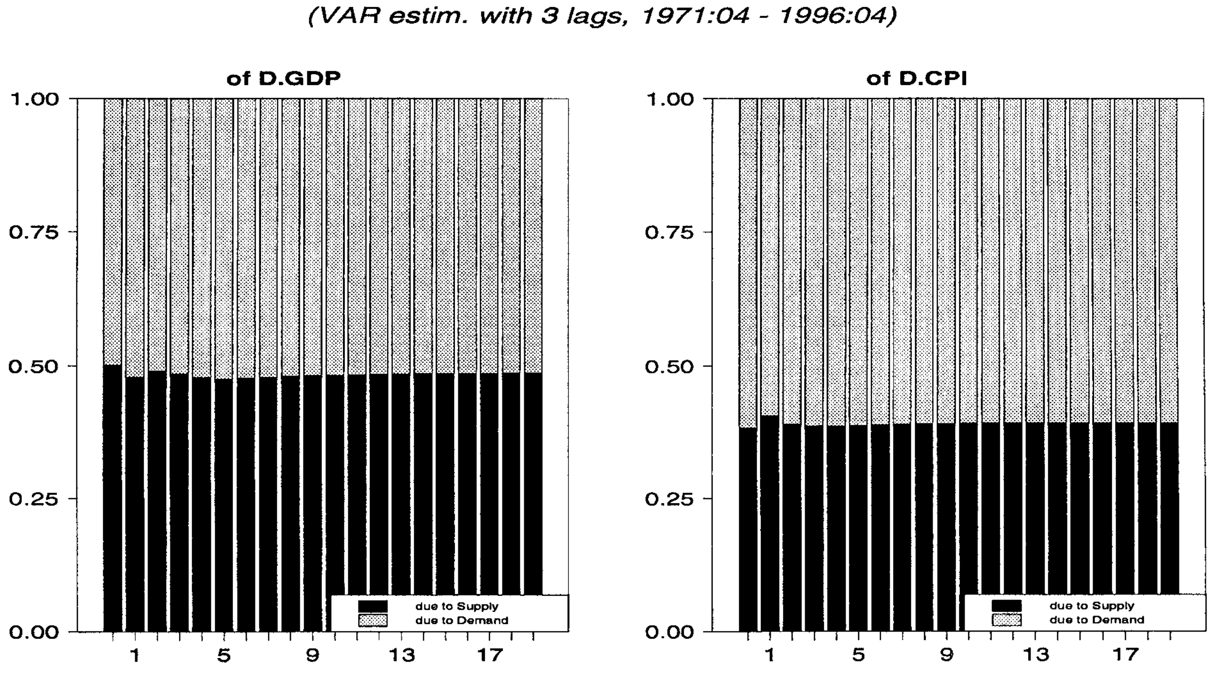
Impulse response functions (bivariate model) – Belgium



Note: Results are those of the bivariate model (output and inflation). The solid lines indicate the estimated and accumulated response to the respective first period structural unit shock, dashed lines above and below are the upper and lower two standard deviation bounds computed from a simulation as described in Appendix A.

Figure 15

Variance decompositions (bivariate model) – Belgium

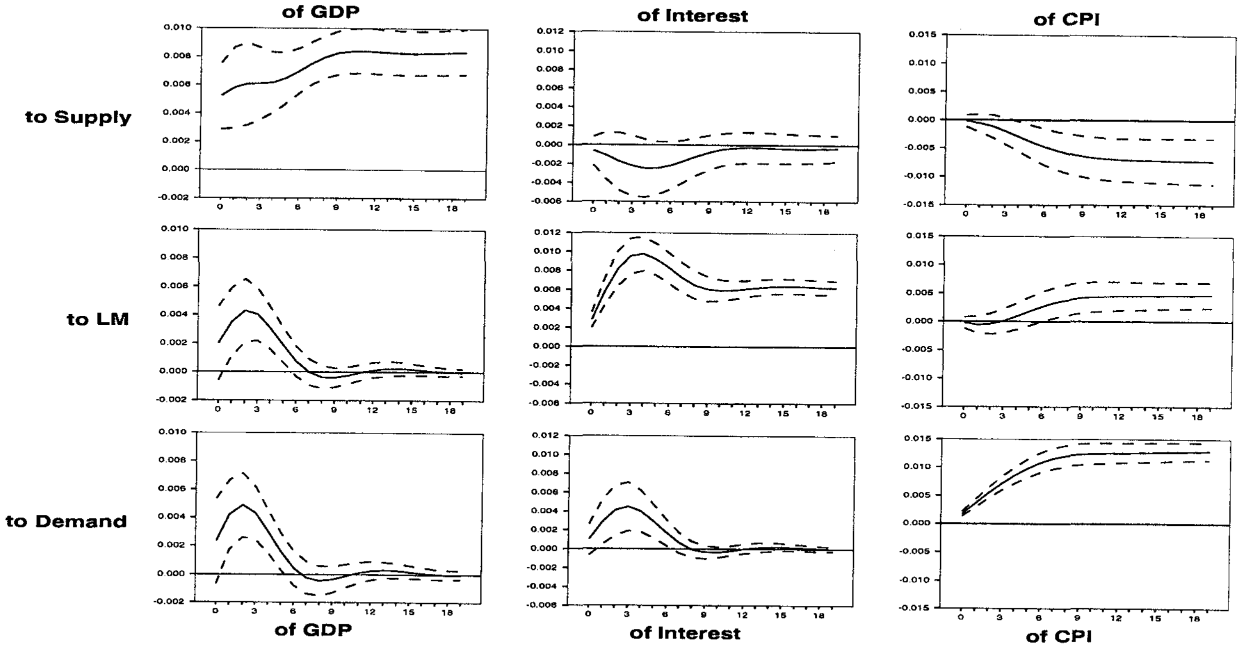


Note: Results are those of the bivariate model (output and inflation). The heights of the respective bars indicate the relative contribution of a specific structural shock to the forecast error variance of the respective series.

Figure 16

Impulse response functions (trivariate model) – Belgium

(VAR estim. with 3 lags, 1971:04 - 1996:04)

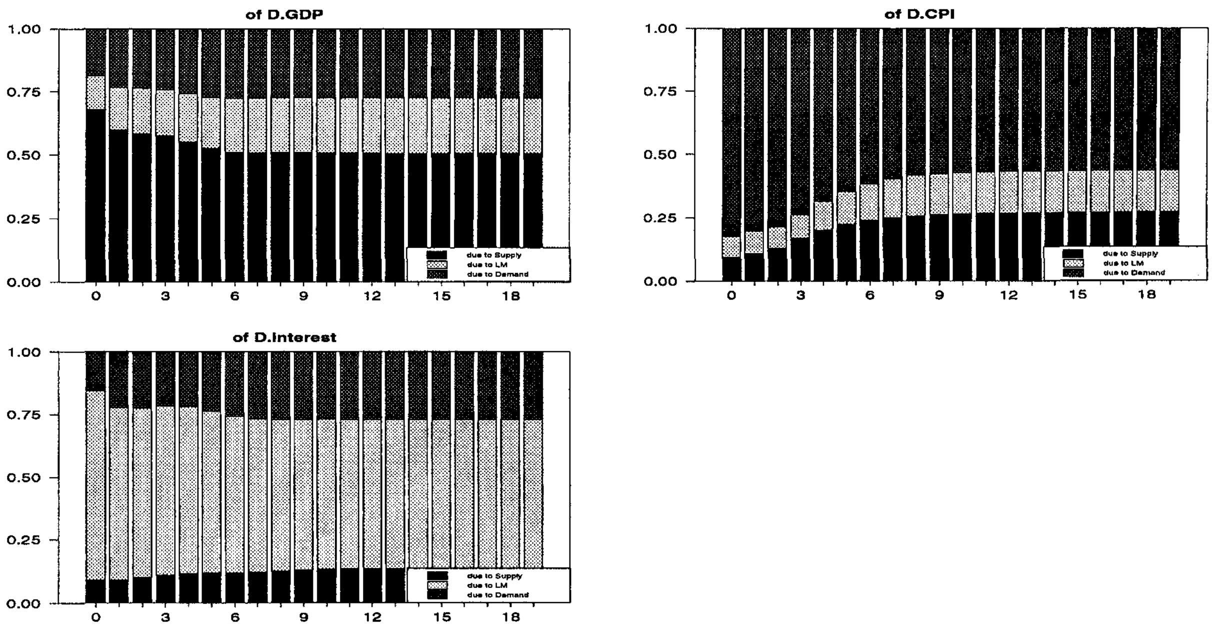


Note: Results are those of the trivariate model (output, interest rate and inflation). The solid lines indicate the estimated and accumulated response to the respective first period structural unit shock, dashed lines above and below are the upper and lower two standard deviation bounds computed from a simulation as described in Appendix A.

Figure 17

Variance decompositions (trivariate model) – Belgium

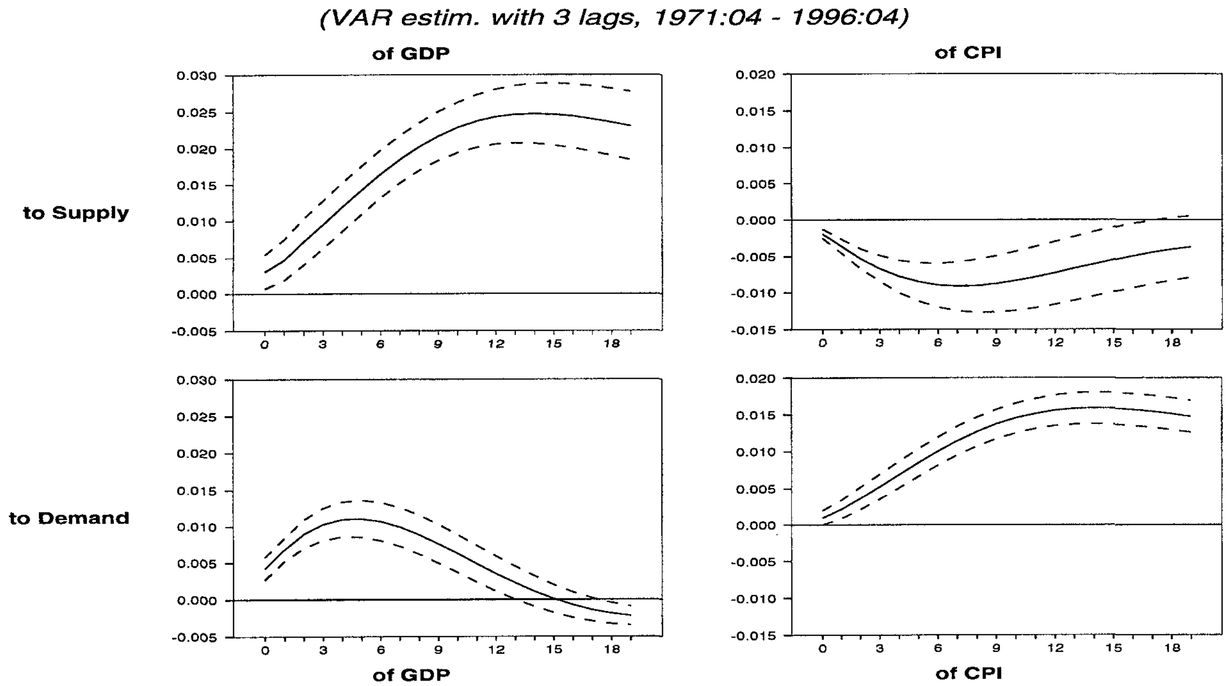
(VAR estim. with 3 lags, 1971:04 - 1996:04)



Note: Results are those of the trivariate model (output, interest rate and inflation). The heights of the respective bars indicate the relative contribution of a specific structural shock to the forecast error variance of the respective series.

Figure 18

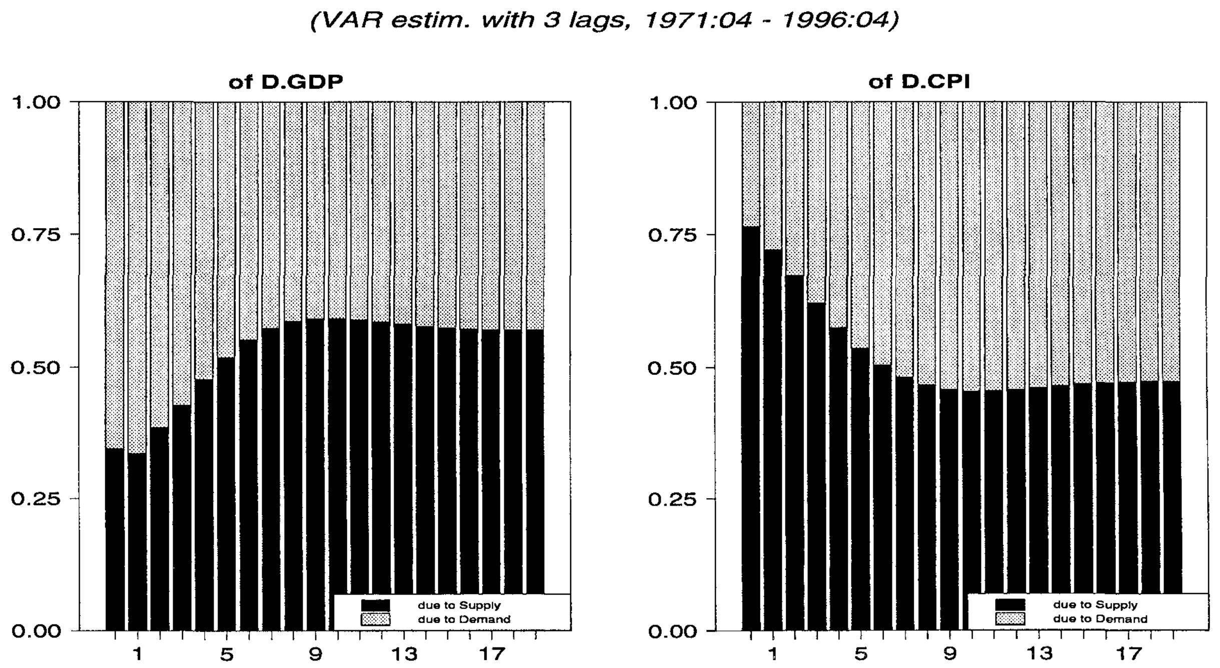
Impulse response functions (bivariate model) – Finland



Note: Results are those of the bivariate model (output and inflation). The solid lines indicate the estimated and accumulated response to the respective first period structural unit shock, dashed lines above and below are the upper and lower two standard deviation bounds computed from a simulation as described in Appendix A.

Figure 19

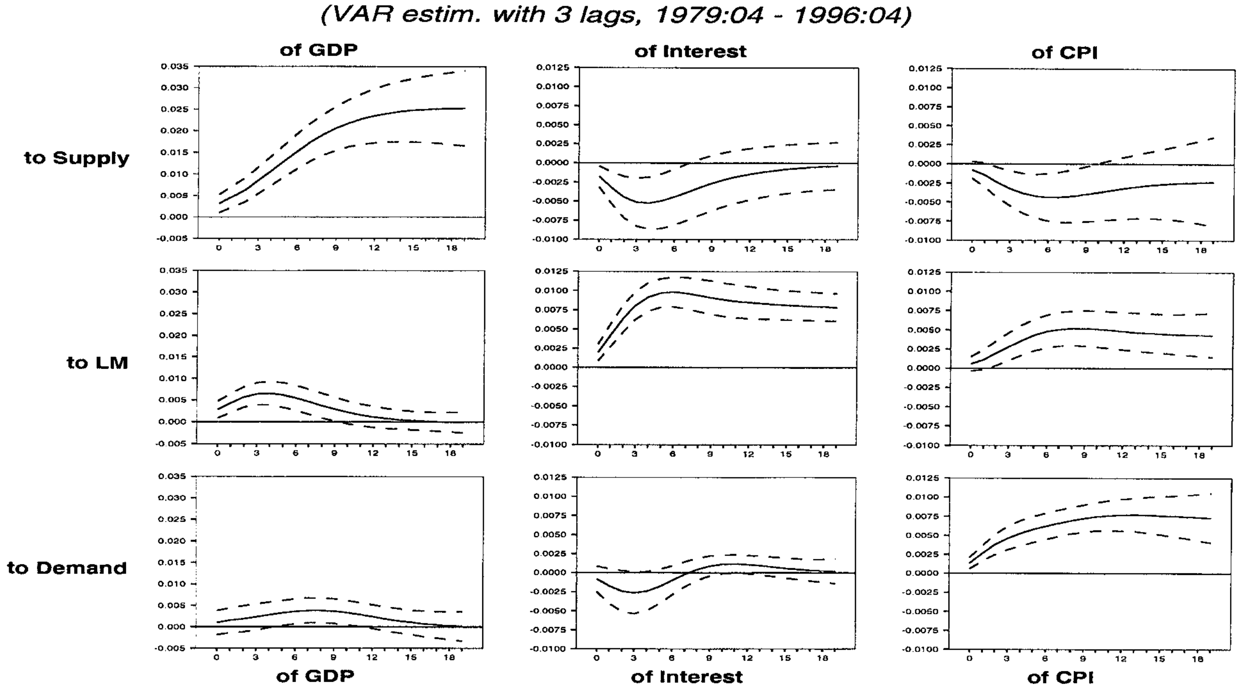
Variance decompositions (bivariate model) – Finland



Note: Results are those of the bivariate model (output and inflation). The heights of the respective bars indicate the relative contribution of a specific structural shock to the forecast error variance of the respective series.

Figure 20

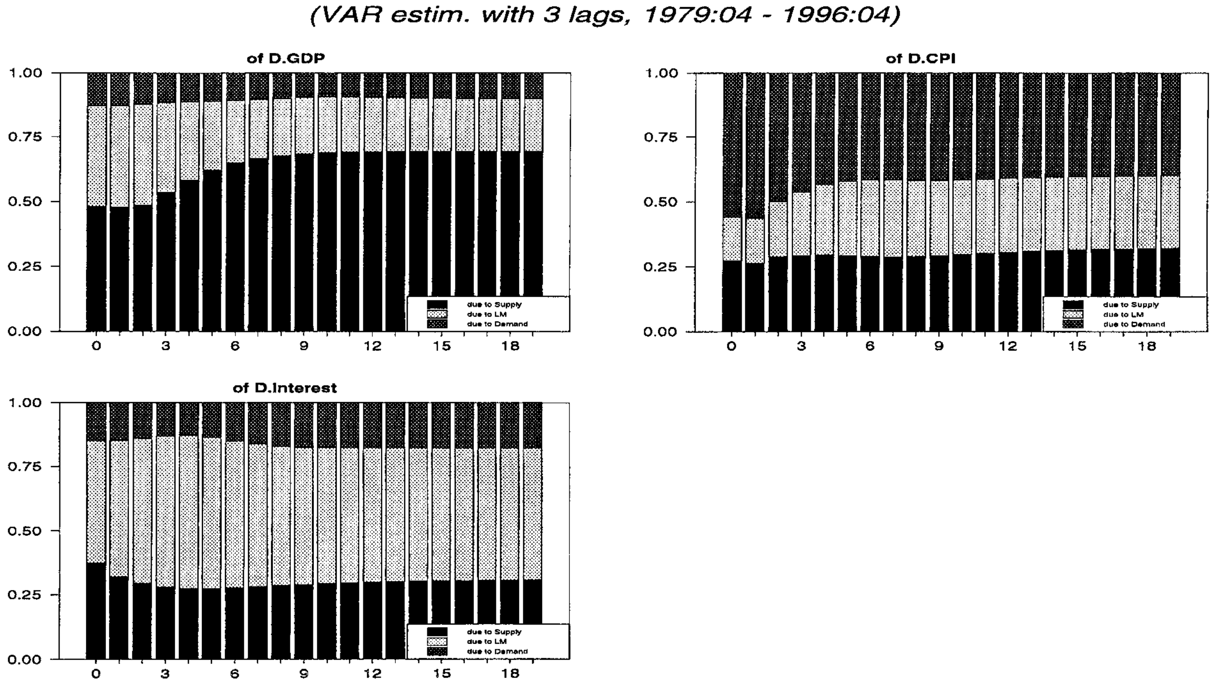
Impulse response functions (trivariate model) – Finland



Note: Results are those of the trivariate model (output, interest rate and inflation). The solid lines indicate the estimated and accumulated response to the respective first period structural unit shock, dashed lines above and below are the upper and lower two standard deviation bounds computed from a simulation as described in Appendix A.

Figure 21

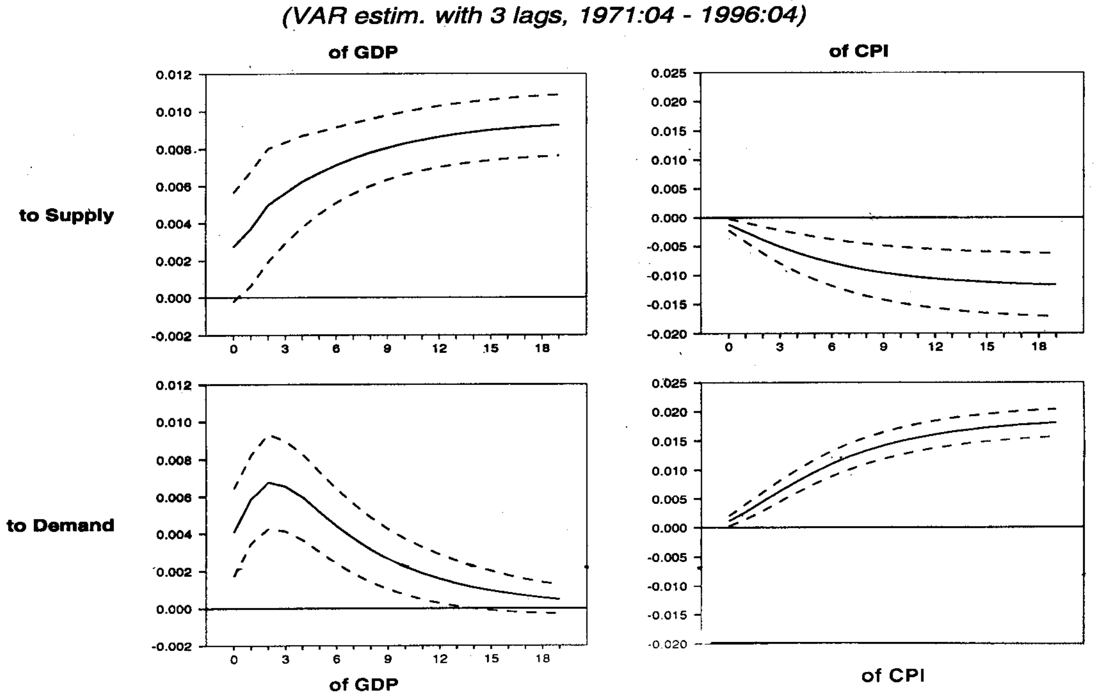
Variance decompositions (trivariate model) – Finland



Note: Results are those of the trivariate model (output, interest rate and inflation). The heights of the respective bars indicate the relative contribution of a specific structural shock to the forecast error variance of the respective series.

Figure 22

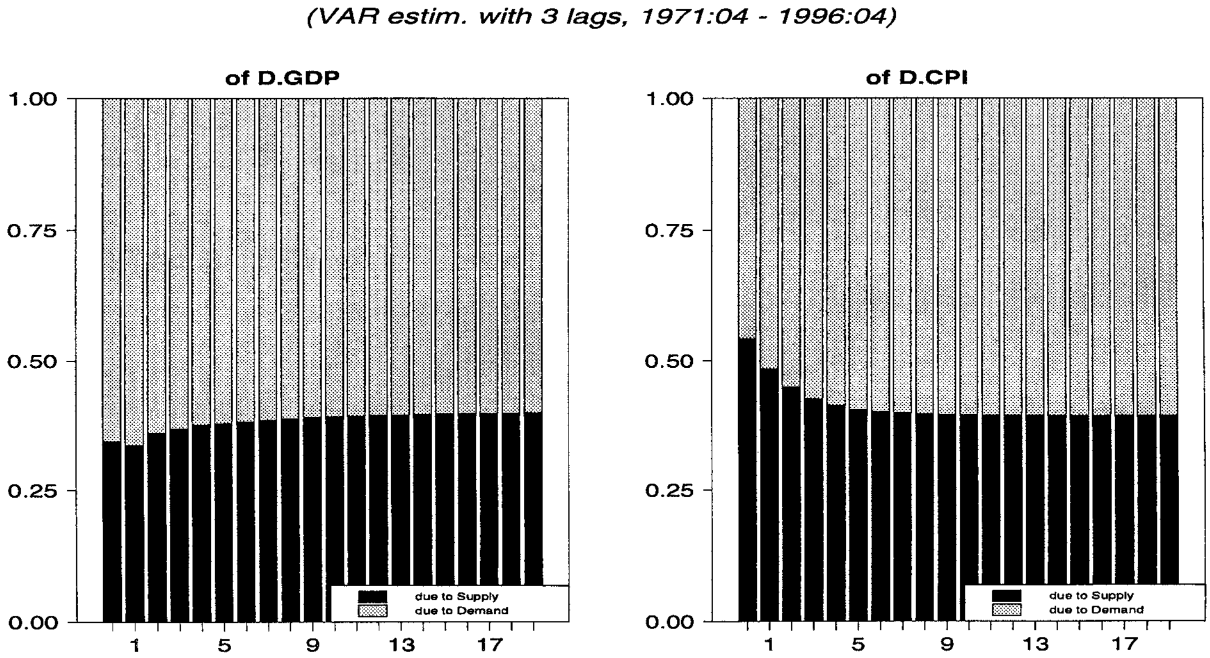
Impulse response functions (bivariate model) – France



Note: Results are those of the bivariate model (output and inflation). The solid lines indicate the estimated and accumulated response to the respective first period structural unit shock, dashed lines above and below are the upper and lower two standard deviation bounds computed from a simulation as described in Appendix A.

Figure 23

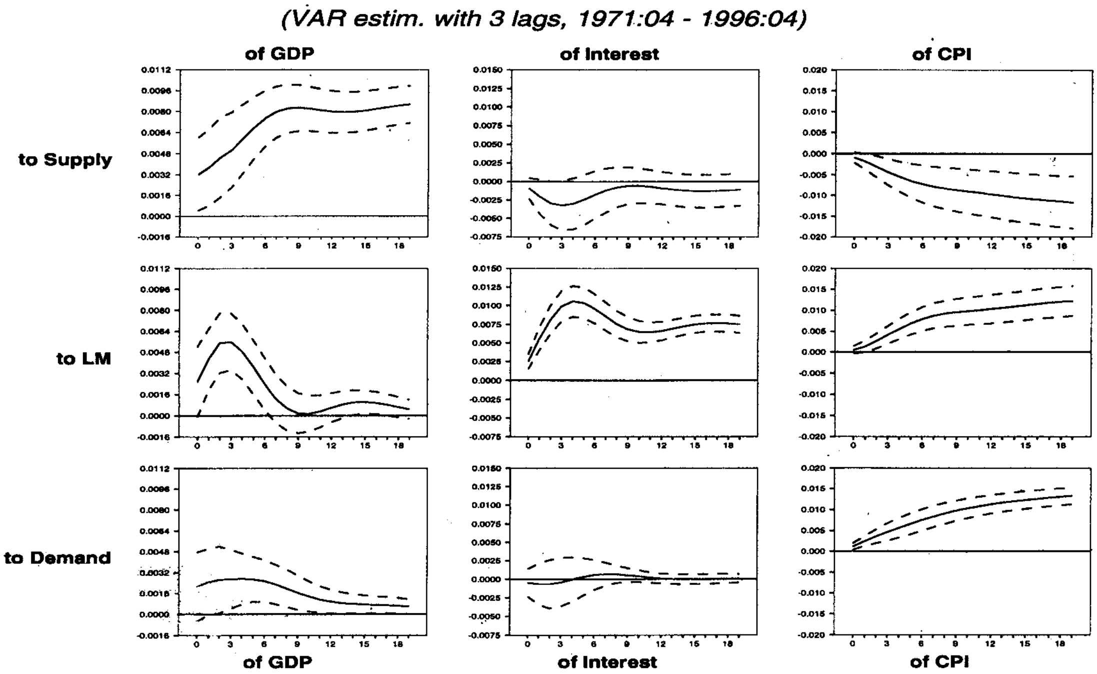
Variance decompositions (bivariate model) – France



Note: Results are those of the bivariate model (output and inflation). The heights of the respective bars indicate the relative contribution of a specific structural shock to the forecast error variance of the respective series.

Figure 24

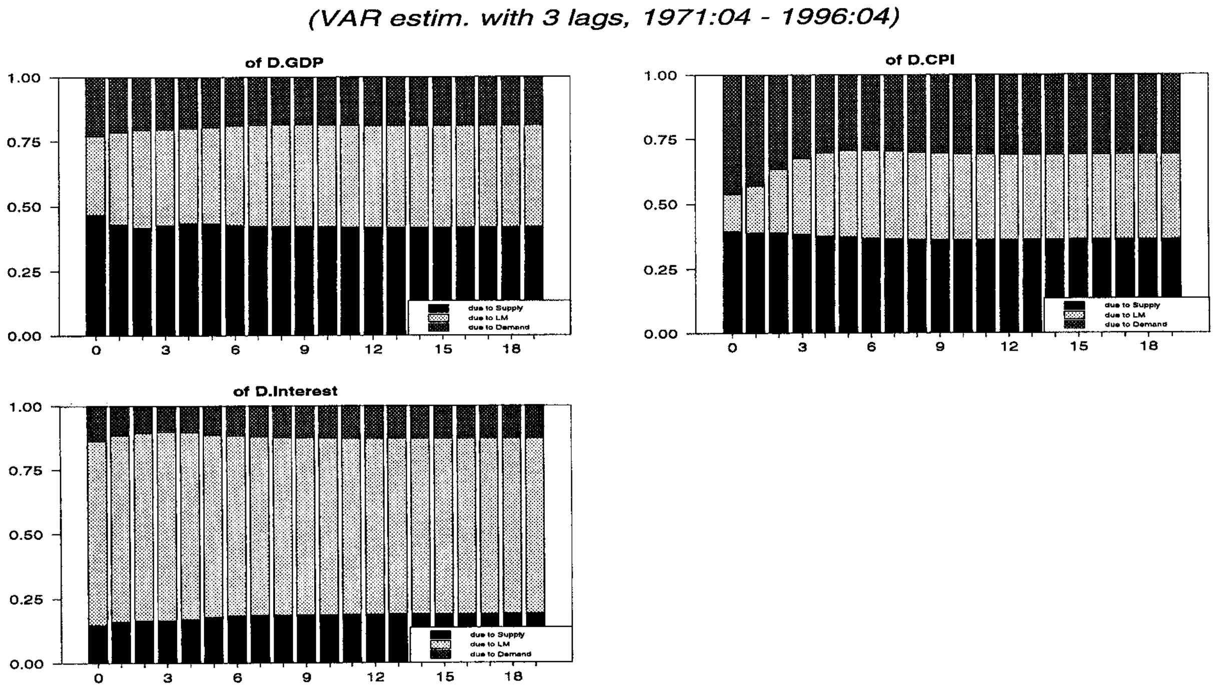
Impulse response functions (trivariate model) – France



Note: Results are those of the trivariate model (output, interest rate and inflation). The solid lines indicate the estimated and accumulated response to the respective first period structural unit shock, dashed lines above and below are the upper and lower two standard deviation bounds computed from a simulation as described in Appendix A.

Figure 25

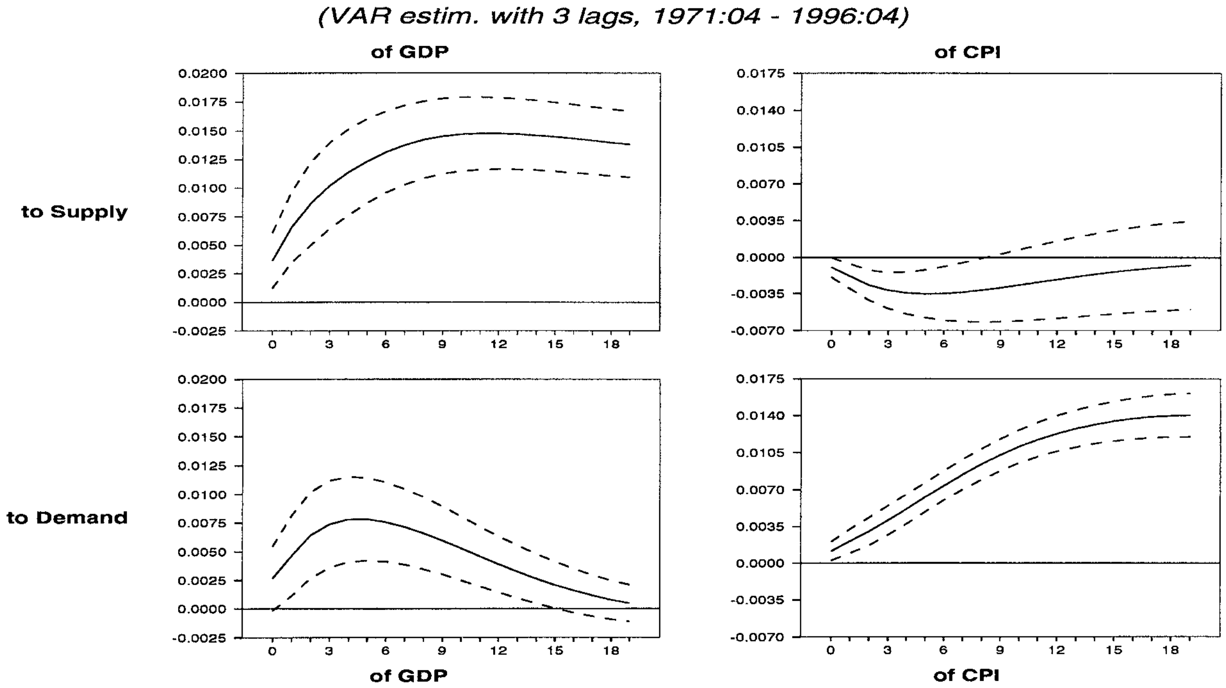
Variance decompositions (trivariate model) – France



Note: Results are those of the trivariate model (output, interest rate and inflation). The heights of the respective bars indicate the relative contribution of a specific structural shock to the forecast error variance of the respective series.

Figure 26

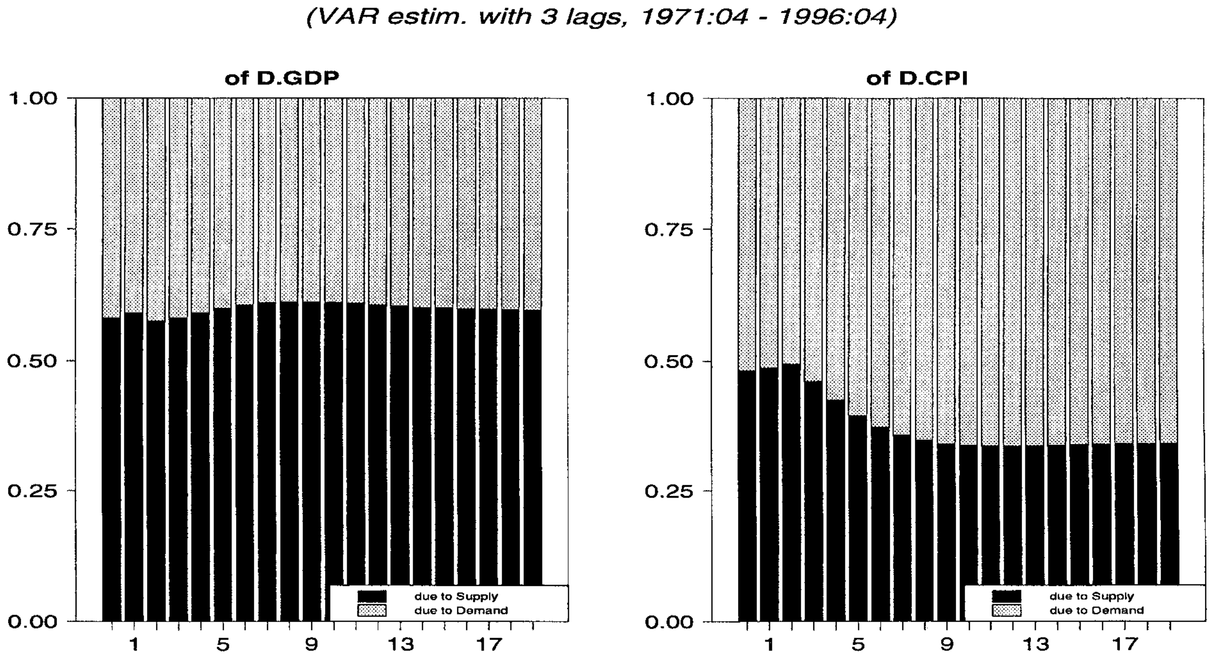
Impulse response functions (bivariate model) – Germany



Note: Results are those of the bivariate model (output and inflation). The solid lines indicate the estimated and accumulated response to the respective first period structural unit shock, dashed lines above and below are the upper and lower two standard deviation bounds computed from a simulation as described in Appendix A.

Figure 27

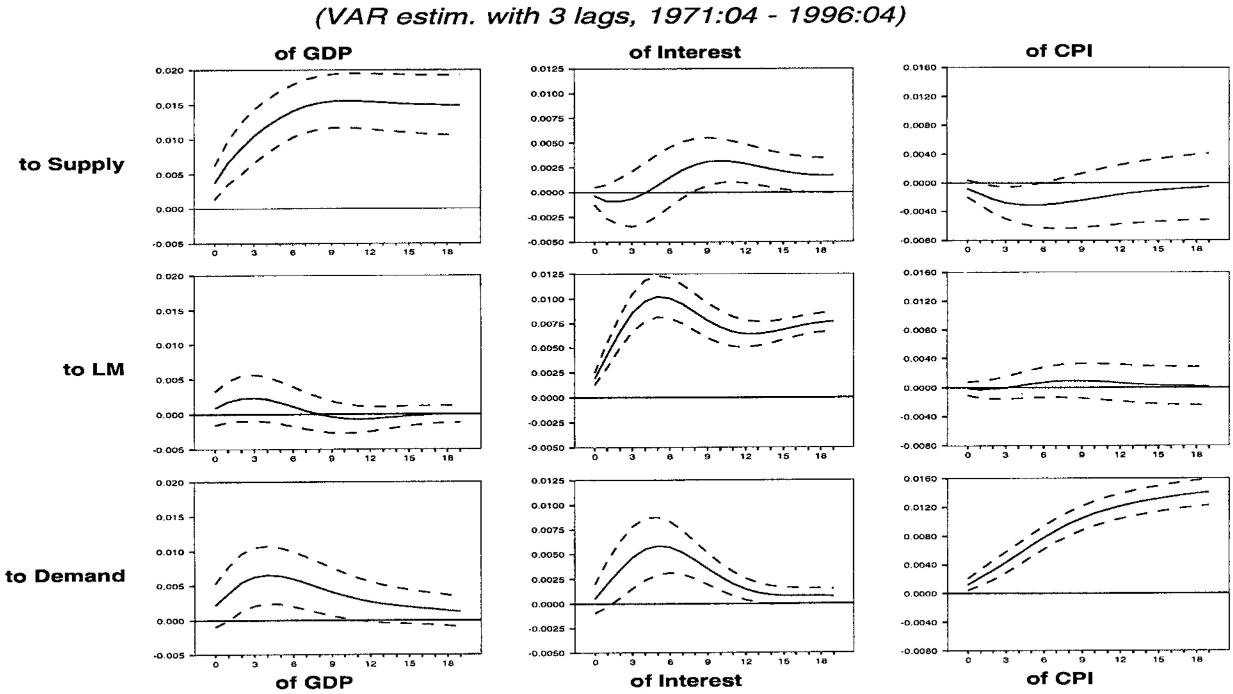
Variance decompositions (bivariate model) – Germany



Note: Results are those of the bivariate model (output and inflation). The heights of the respective bars indicate the relative contribution of a specific structural shock to the forecast error variance of the respective series.

Figure 28

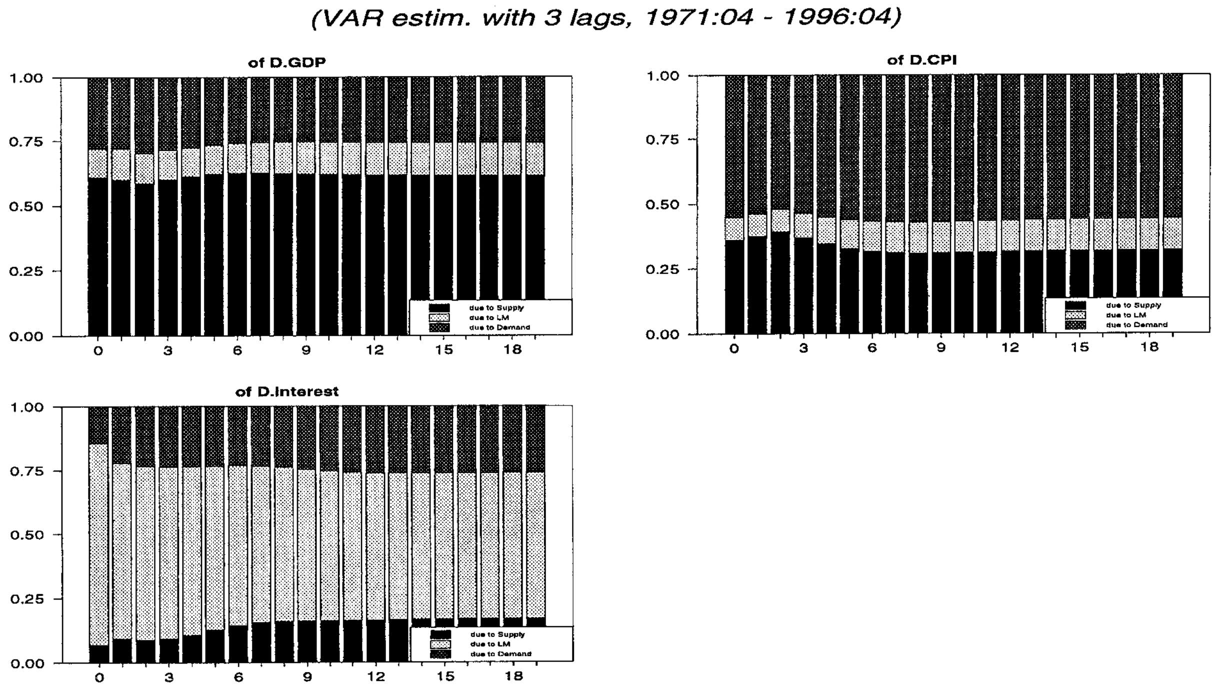
Impulse response functions (trivariate model) – Germany



Note: Results are those of the trivariate model (output, interest rate and inflation). The solid lines indicate the estimated and accumulated response to the respective first period structural unit shock, dashed lines above and below are the upper and lower two standard deviation bounds computed from a simulation as described in Appendix A.

Figure 29

Variance decompositions (trivariate model) – Germany

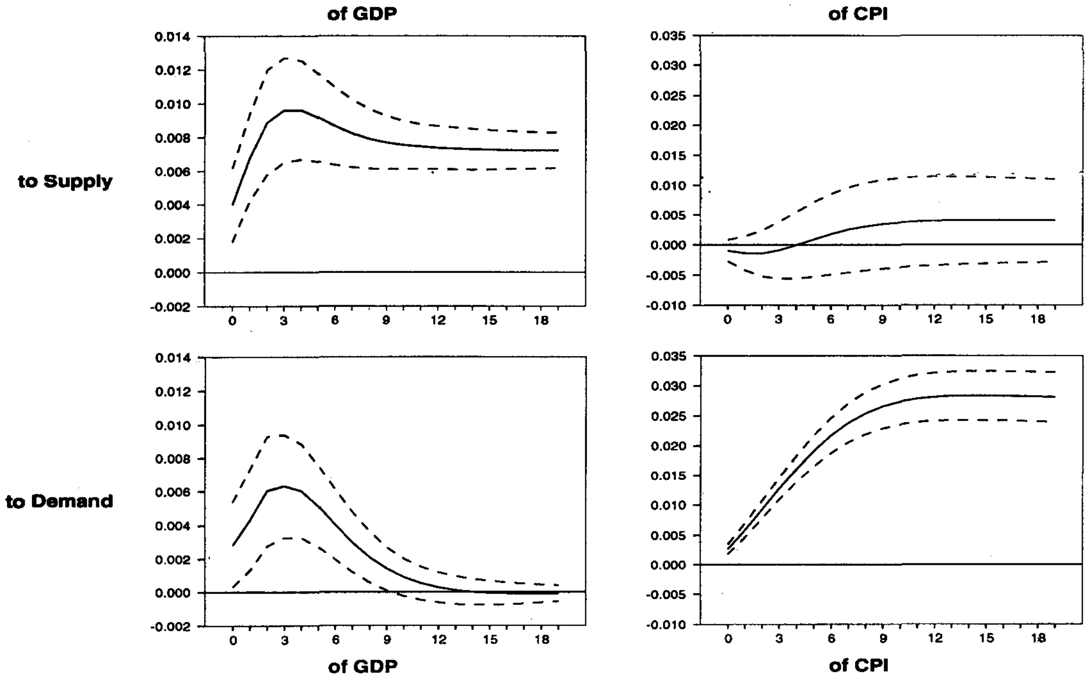


Note: Results are those of the trivariate model (output, interest rate and inflation). The heights of the respective bars indicate the relative contribution of a specific structural shock to the forecast error variance of the respective series.

Figure 30

Impulse response functions (bivariate model) – Italy

(VAR estim. with 3 lags, 1971:04 - 1996:04)

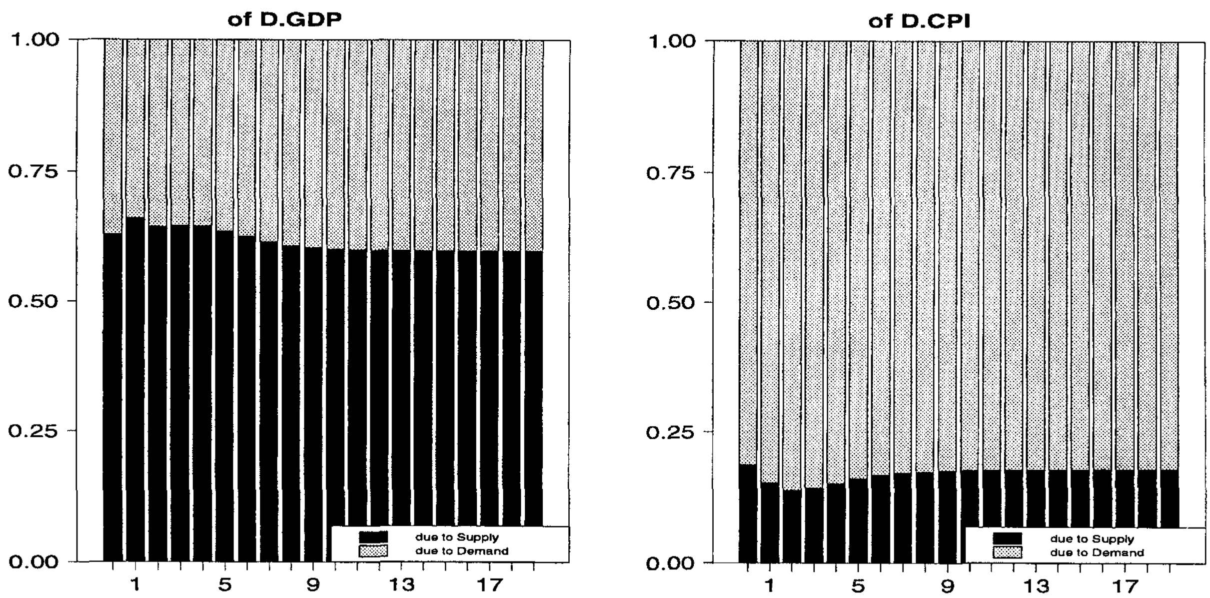


Note: Results are those of the bivariate model (output and inflation). The solid lines indicate the estimated and accumulated response to the respective first period structural unit shock, dashed lines above and below are the upper and lower two standard deviation bounds computed from a simulation as described in Appendix A.

Figure 31

Variance decompositions (bivariate model) – Italy

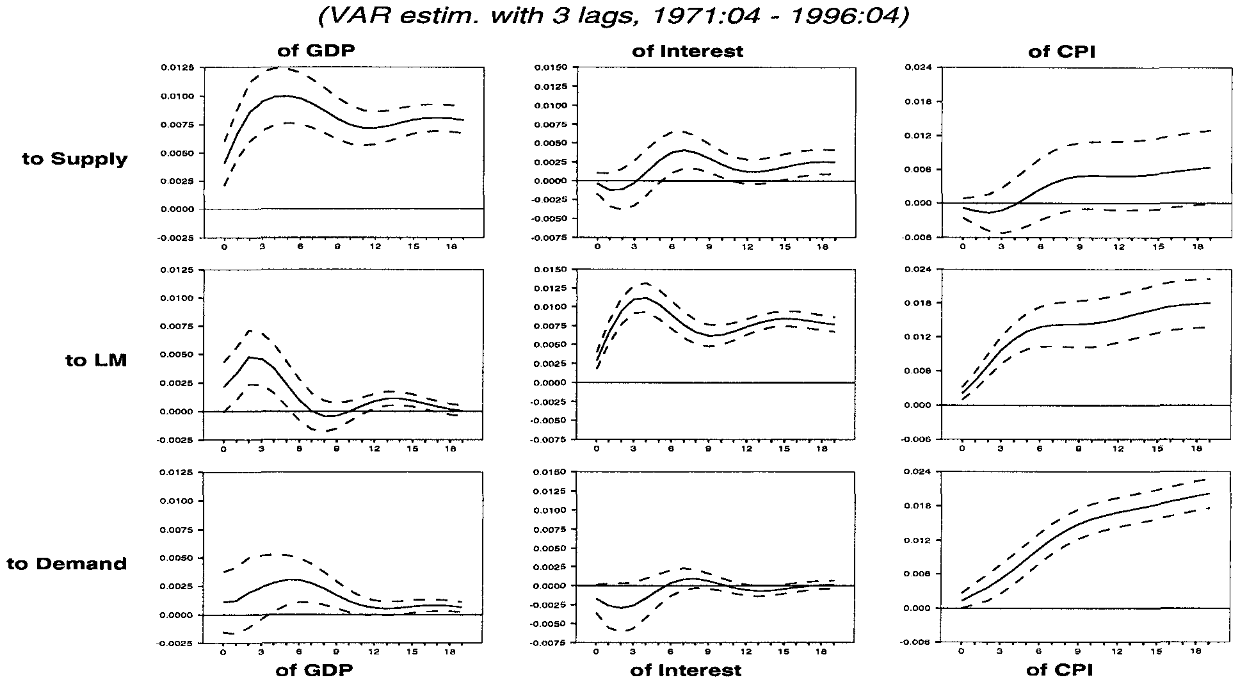
(VAR estim. with 3 lags, 1971:04 - 1996:04)



Note: Results are those of the bivariate model (output and inflation). The heights of the respective bars indicate the relative contribution of a specific structural shock to the forecast error variance of the respective series.

Figure 32

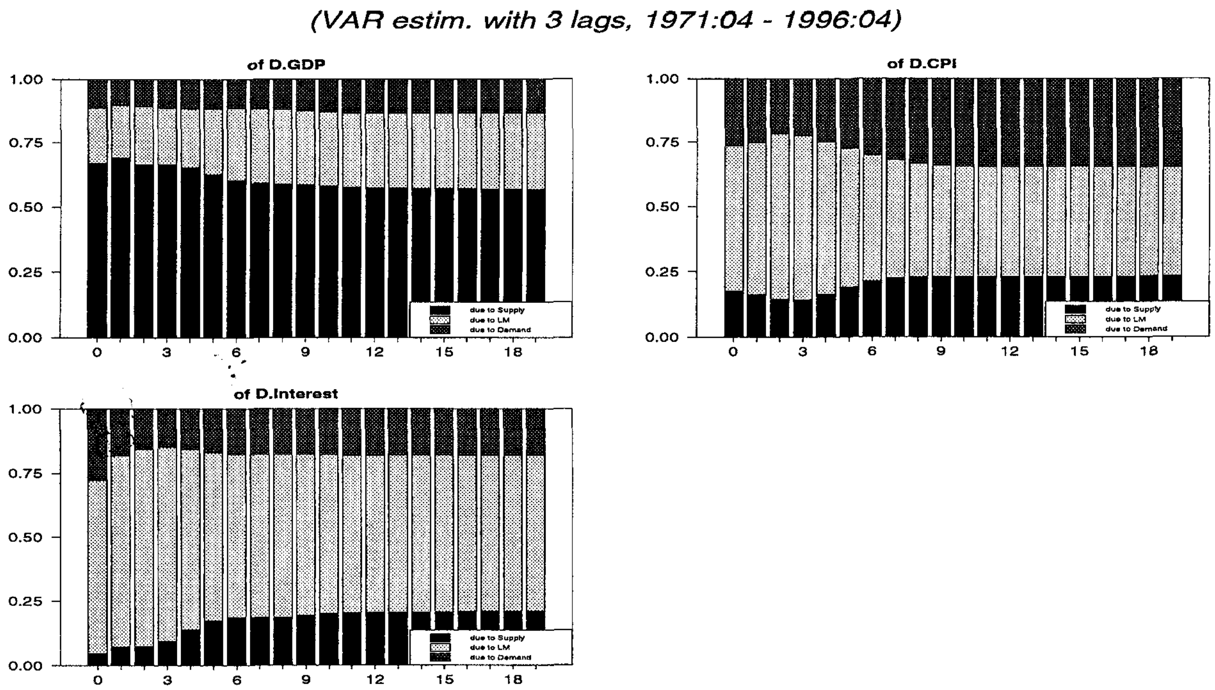
Impulse response functions (trivariate model) – Italy



Note: Results are those of the trivariate model (output, interest rate and inflation). The solid lines indicate the estimated and accumulated response to the respective first period structural unit shock, dashed lines above and below are the upper and lower two standard deviation bounds computed from a simulation as described in Appendix A.

Figure 33

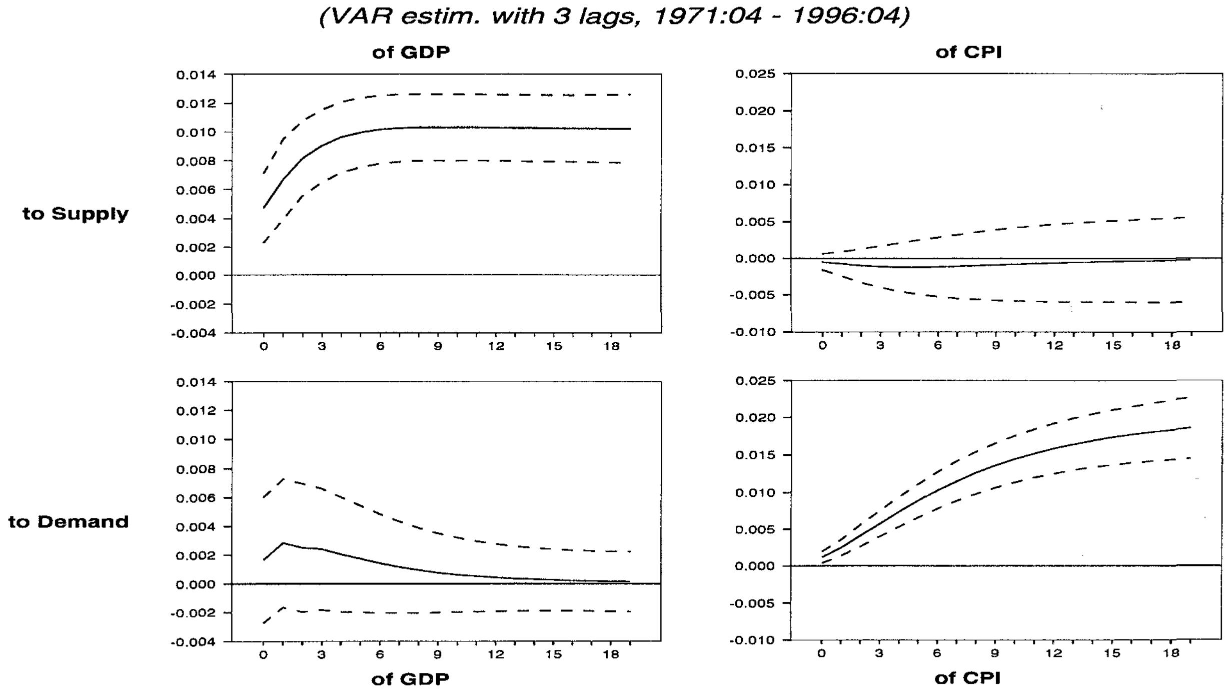
Variance decompositions (trivariate model) – Italy



Note: Results are those of the trivariate model (output, interest rate and inflation). The heights of the respective bars indicate the relative contribution of a specific structural shock to the forecast error variance of the respective series.

Figure 34

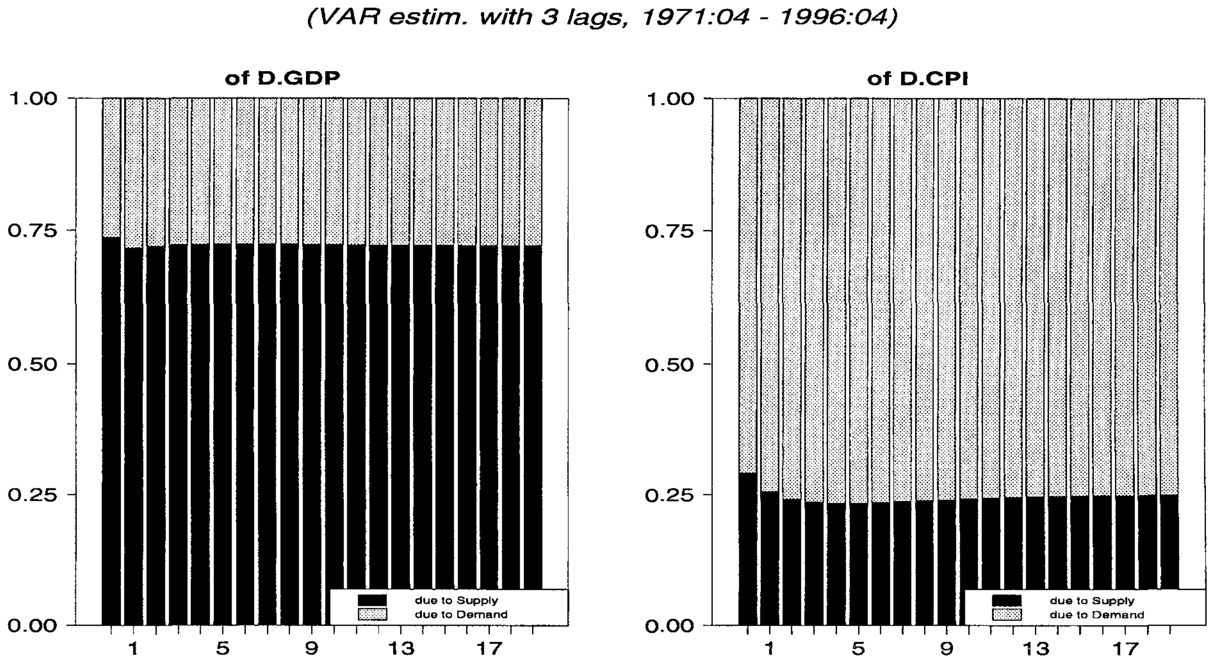
Impulse response functions (bivariate model) – Netherlands



Note: Results are those of the bivariate model (output and inflation). The solid lines indicate the estimated and accumulated response to the respective first period structural unit shock, dashed lines above and below are the upper and lower two standard deviation bounds computed from a simulation as described in Appendix A.

Figure 35

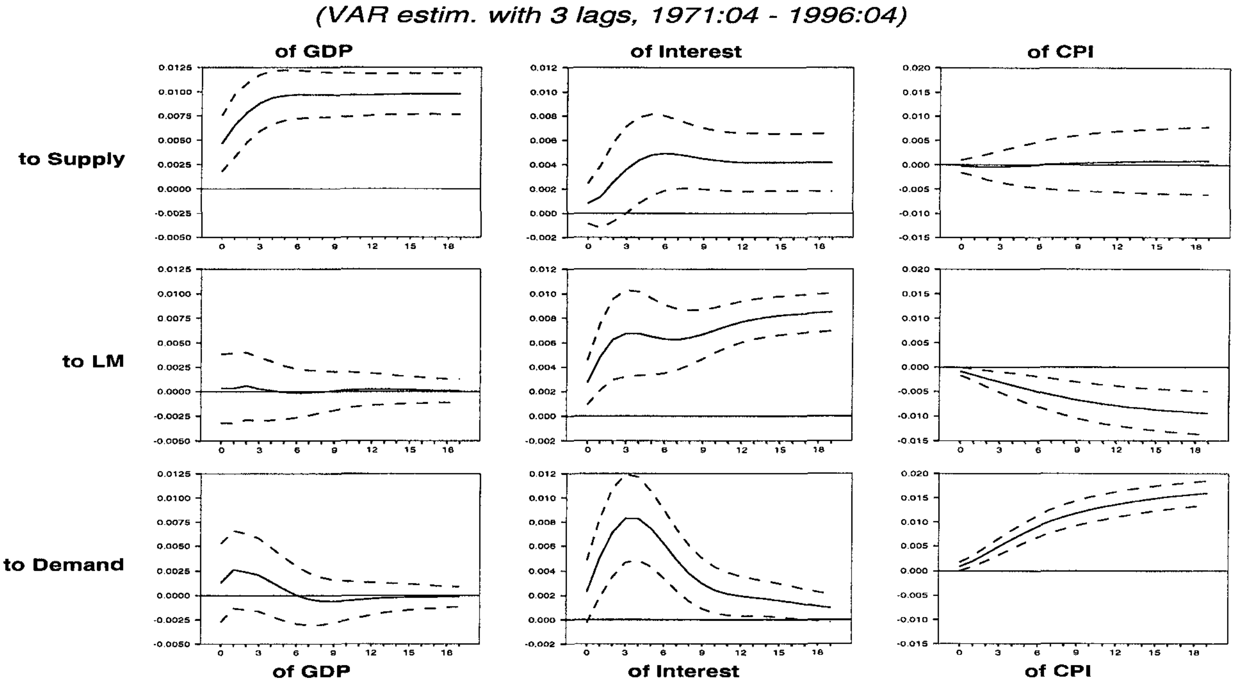
Variance decompositions (bivariate model) – Netherlands



Note: Results are those of the bivariate model (output and inflation). The heights of the respective bars indicate the relative contribution of a specific structural shock to the forecast error variance of the respective series.

Figure 36

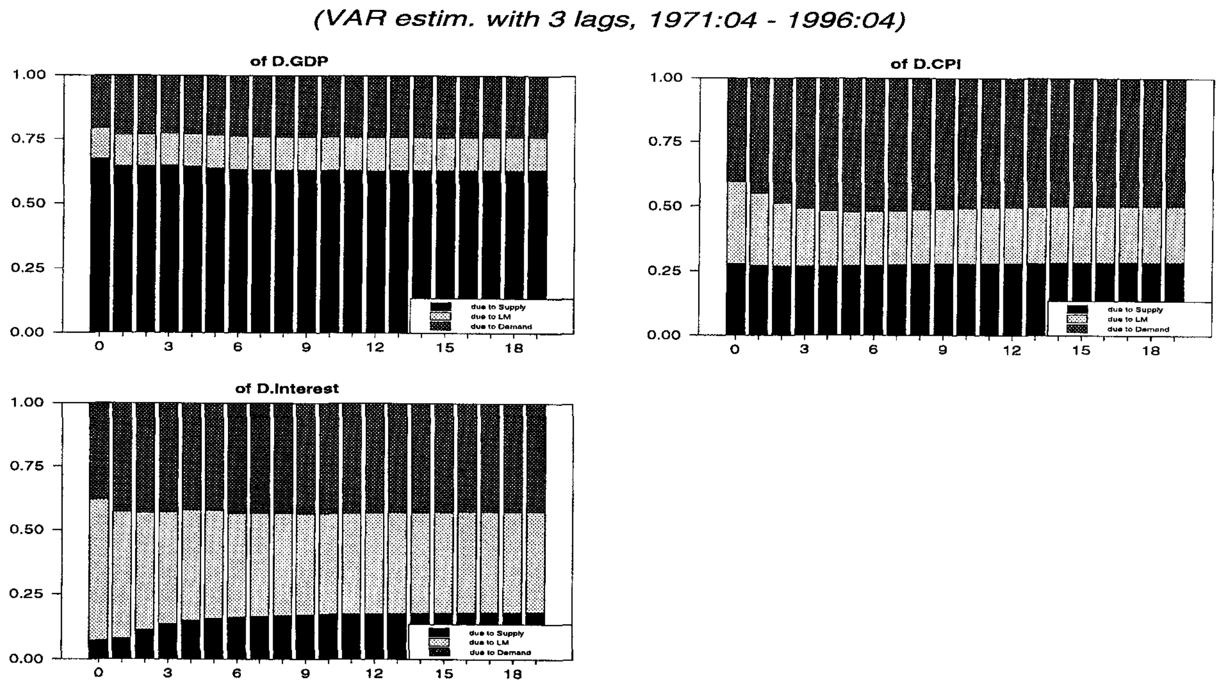
Impulse response functions (trivariate model) – Netherlands



Note: Results are those of the trivariate model (output, interest rate and inflation). The solid lines indicate the estimated and accumulated response to the respective first period structural unit shock, dashed lines above and below are the upper and lower two standard deviation bounds computed from a simulation as described in Appendix A.

Figure 37

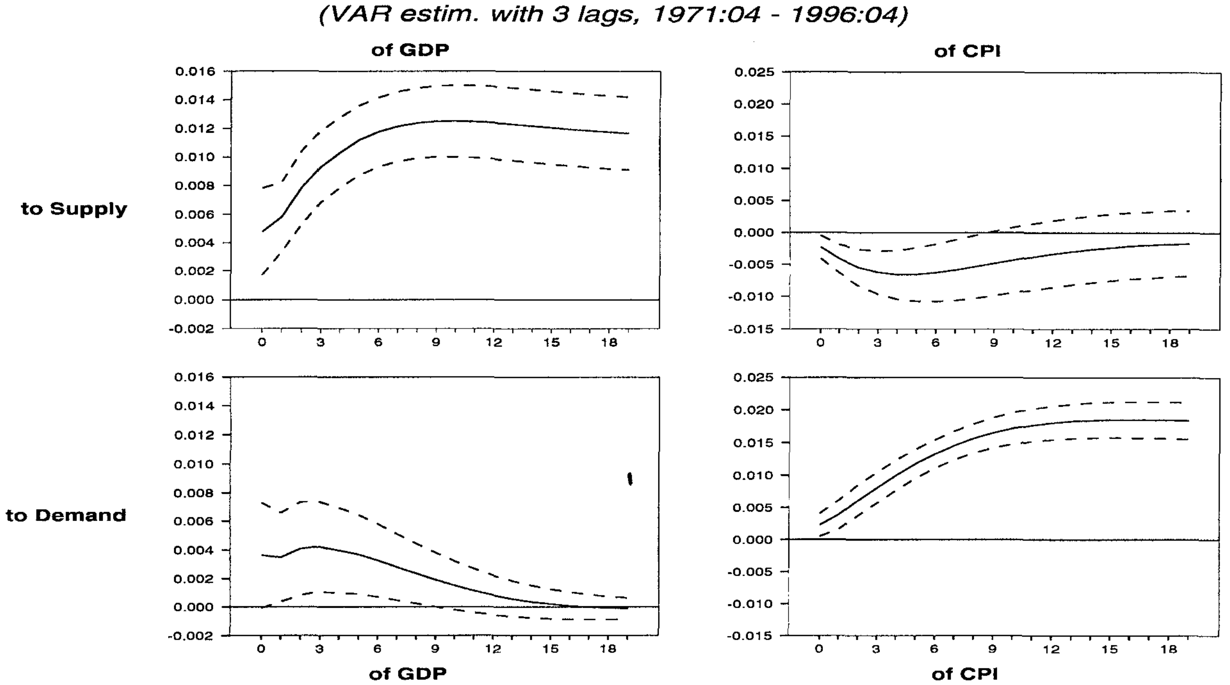
Variance decompositions (trivariate model) – Netherlands



Note: Results are those of the trivariate model (output, interest rate and inflation). The heights of the respective bars indicate the relative contribution of a specific structural shock to the forecast error variance of the respective series.

Figure 38

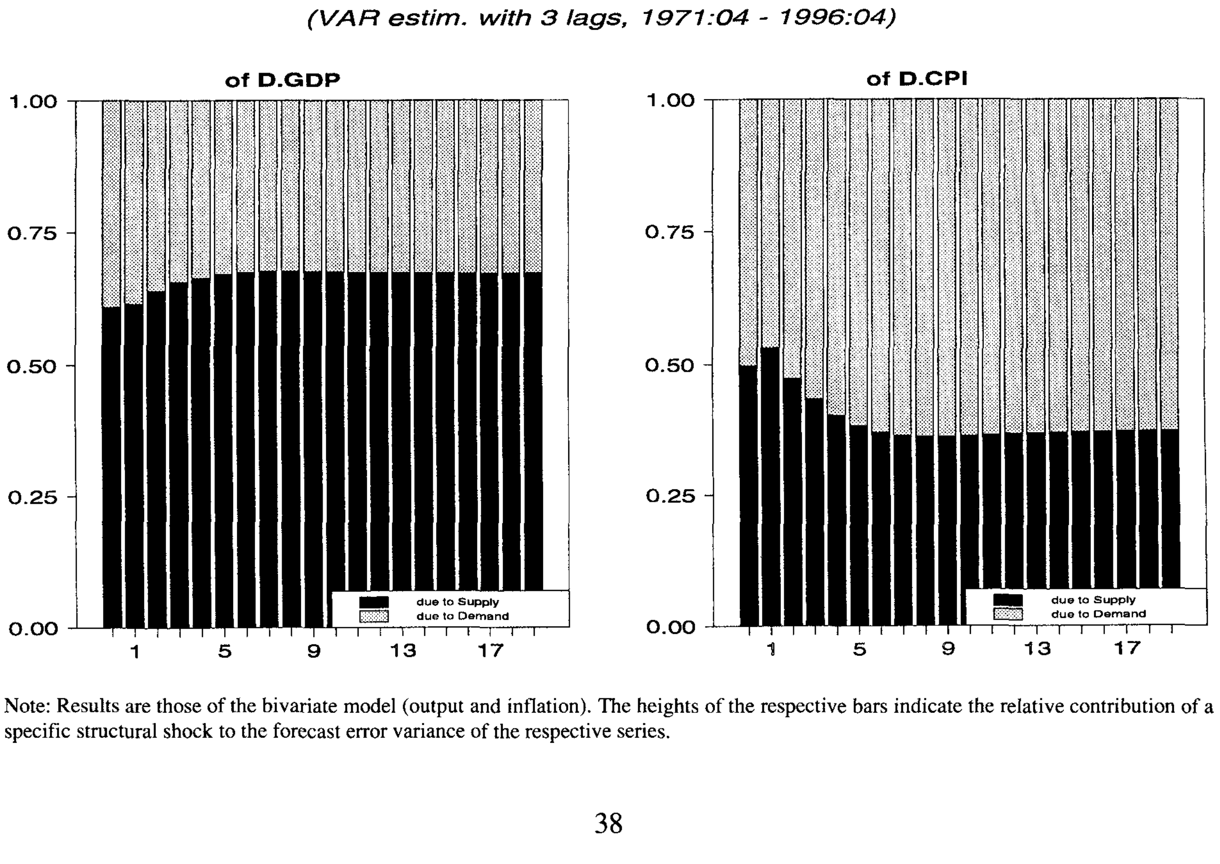
Impulse response functions (bivariate model) – Sweden



Note: Results are those of the bivariate model (output and inflation). The solid lines indicate the estimated and accumulated response to the respective first period structural unit shock, dashed lines above and below are the upper and lower two standard deviation bounds computed from a simulation as described in Appendix A.

Figure 39

Variance decompositions (bivariate model) – Sweden

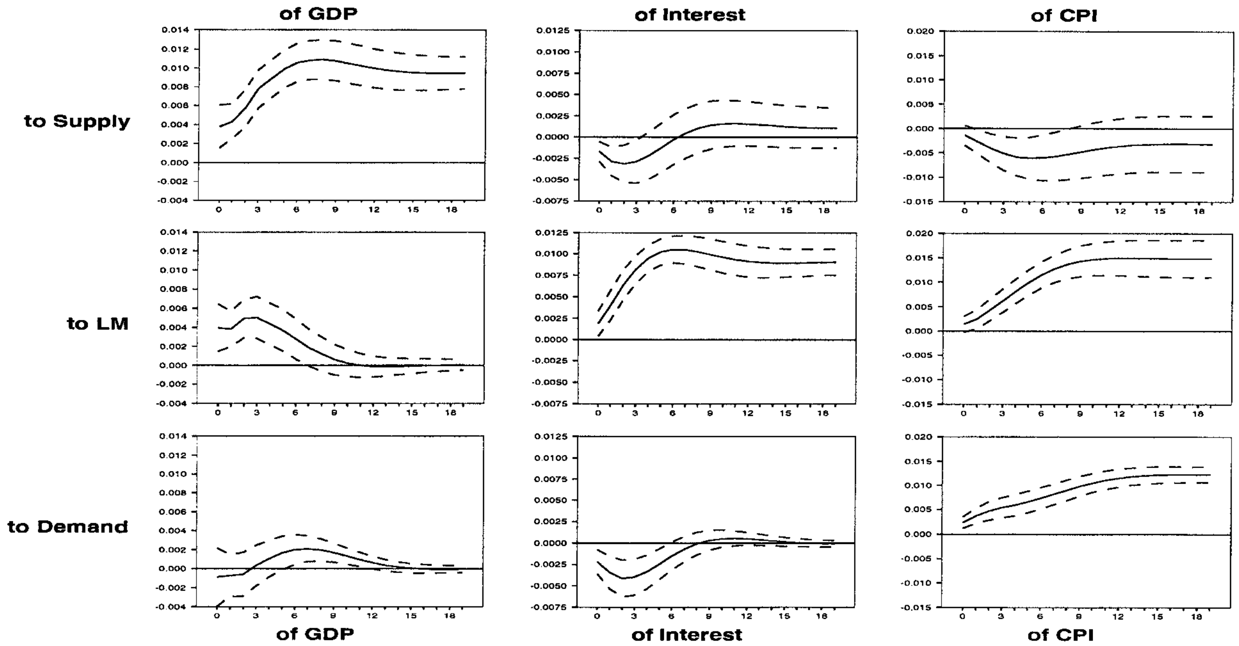


Note: Results are those of the bivariate model (output and inflation). The heights of the respective bars indicate the relative contribution of a specific structural shock to the forecast error variance of the respective series.

Figure 40

Impulse response functions (trivariate model) – Sweden

(VAR estim. with 3 lags, 1971:04 - 1996:04)

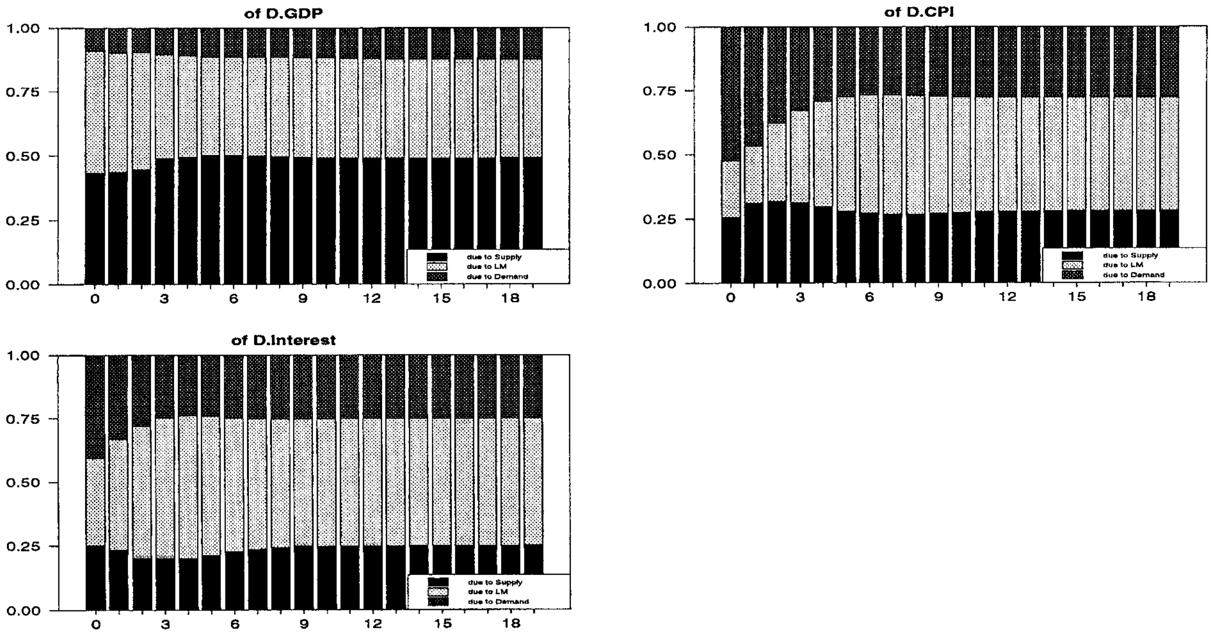


Note: Results are those of the trivariate model (output, interest rate and inflation). The solid lines indicate the estimated and accumulated response to the respective first period structural unit shock, dashed lines above and below are the upper and lower two standard deviation bounds computed from a simulation as described in Appendix A.

Figure 41

Variance decompositions (trivariate model) – Sweden

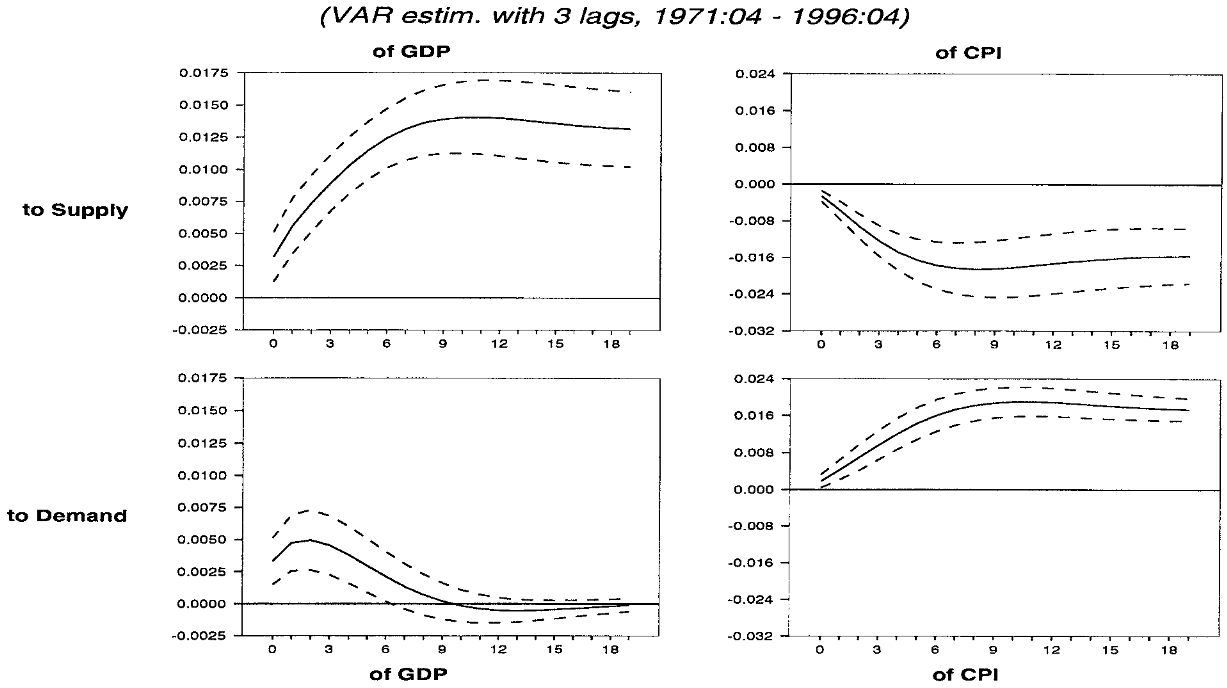
(VAR estim. with 3 lags, 1971:04 - 1996:04)



Note: Results are those of the trivariate model (output, interest rate and inflation). The heights of the respective bars indicate the relative contribution of a specific structural shock to the forecast error variance of the respective series.

Figure 42

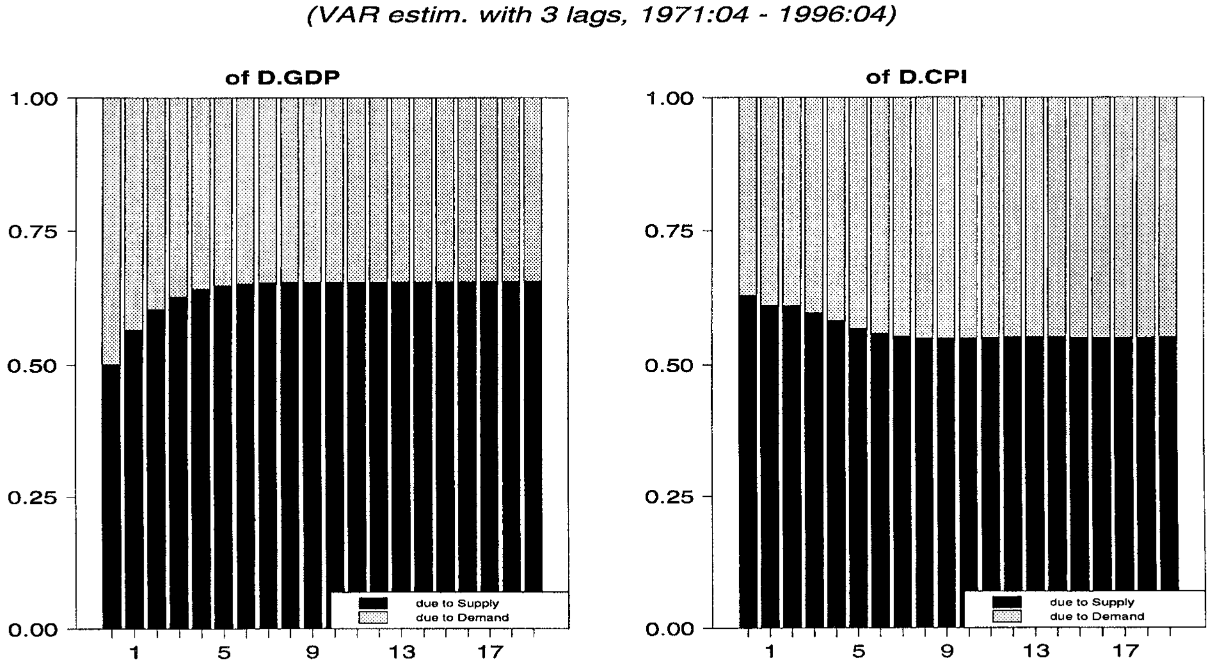
Impulse response functions (bivariate model) – United Kingdom



Note: Results are those of the bivariate model (output and inflation). The solid lines indicate the estimated and accumulated response to the respective first period structural unit shock, dashed lines above and below are the upper and lower two standard deviation bounds computed from a simulation as described in Appendix A.

Figure 43

Variance decompositions (bivariate model) – United Kingdom

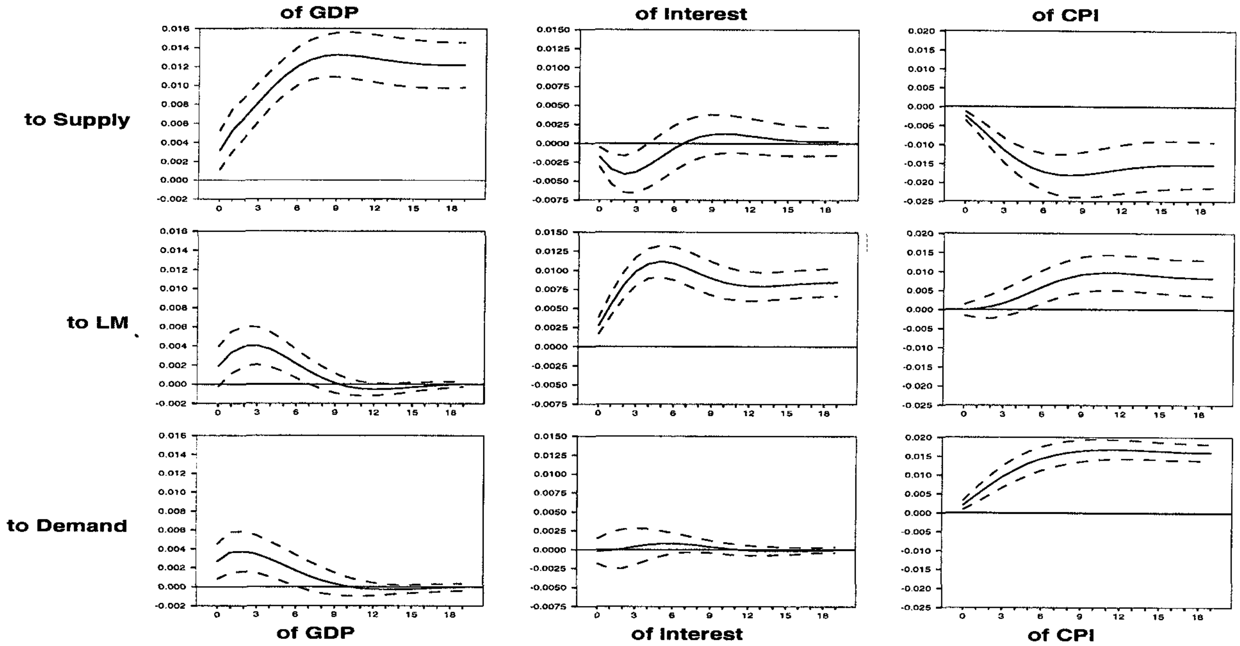


Note: Results are those of the bivariate model (output and inflation). The heights of the respective bars indicate the relative contribution of a specific structural shock to the forecast error variance of the respective series.

Figure 44

Impulse response functions (trivariate model) – United Kingdom

(VAR estim. with 3 lags, 1971:04 - 1996:04)

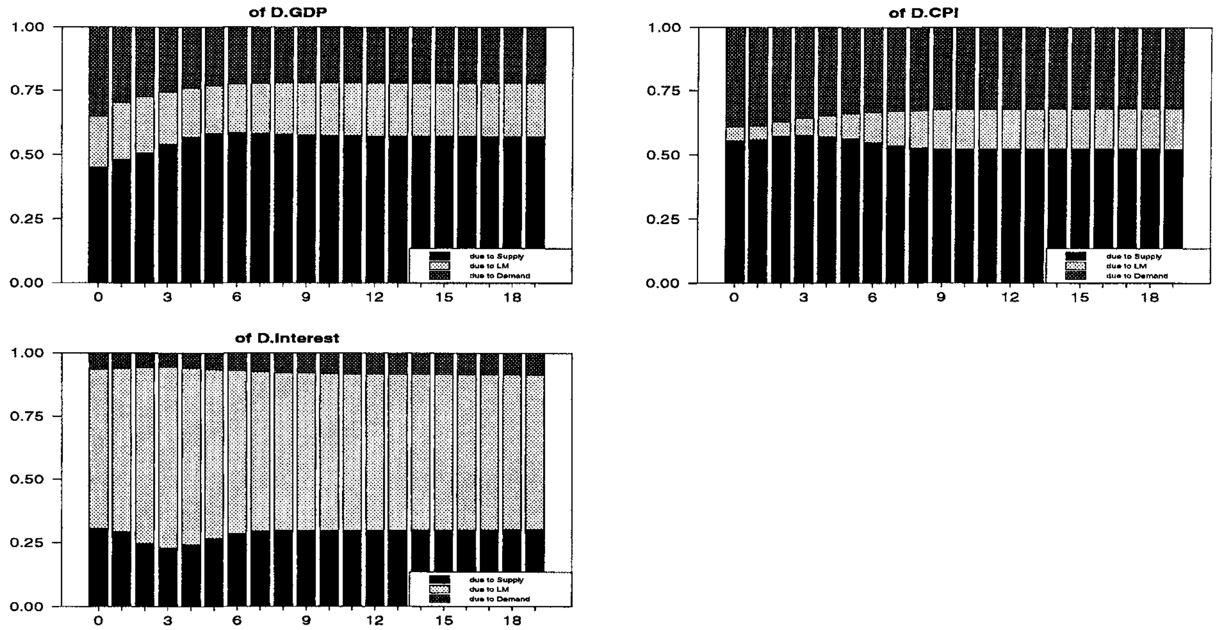


Note: Results are those of the trivariate model (output, interest rate and inflation). The solid lines indicate the estimated and accumulated response to the respective first period structural unit shock, dashed lines above and below are the upper and lower two standard deviation bounds computed from a simulation as described in Appendix A.

Figure 45

Variance decompositions (trivariate model) – United Kingdom

(VAR estim. with 3 lags, 1971:04 - 1996:04)



Note: Results are those of the trivariate model (output, interest rate and inflation). The heights of the respective bars indicate the relative contribution of a specific structural shock to the forecast error variance of the respective series.

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Comments on: "Core inflation in selected European Union countries" by Christine Gartner and Gert D. Wehinger

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by Carsten K. Folkertsma

Introduction

This well-written and interesting paper contains a wealth of information on inflation and its causes in nine EMU countries. Especially interesting is the extension of the bivariate structural VAR-model of Quah and Vahey with the short-term interest rate as an indicator of monetary policy. This extension might be a direct method to identify the purely monetary component of measured inflation. Another nice feature of the study is the fact that the authors decompose measured inflation using the same model specification and sample period for each country. This allows comparisons across countries of the inflation-responses to various shocks.

My comments will be on two topics: first, the measurement of inflation in general and the differences between the approach of Gartner-Wehinger and Quah-Vahey and second on the usefulness of the results for the monetary policy of the ECB.

1. Measuring inflation

It seems a widely accepted fact that there is a mismatch between our measurement of inflation by means of the CPI and our theoretical concept of inflation as the *sustained increase of the general price level.* This mismatch is due to the fact that the CPI is just a weighted average of consumer prices. As such, it is based only on a small subset of prices and it cannot distinguish between transitory and sustained price increases. Moreover, due to the weighting of the CPI some price changes have a greater impact on measured inflation than others.

There have been various proposals to refine the CPI changes by means of smoothing techniques or zero weighting of certain components. However, none of these approaches is fully convincing since they contain an element of arbitrariness and they lack a firm theoretical foundation.

An altogether different approach has been suggested by Quah and Vahey. They proposed a method to identify what they call core inflation by using the theoretical insight that inflation has no long run impact on output. This long run restriction is sufficient to recover the core inflation series by means of a bivariate VAR model explaining output growth and the change of measured inflation. Note, that Quah and Vahey assume in their analysis that inflation is integrated of order one, in order to identify that part of measured inflation that is output-neutral in the long run.

The study of Gartner and Wehinger deviates from Quah and Vahey's approach in an important aspect. From their preliminary data analysis, they conclude that inflation is trend-stationary and not integrated of order one.[1] Consequently, Gartner and Wehinger then proceed to apply Quah and Vahey's identification scheme to inflation and not to its first difference. As a result Gartner and

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[1] Quah and Vahey used monthly data for the sample period 1969:3–1994:3. Their inflation series was based on the UK Retail Price Index.

Wehinger identify not only the component of measured inflation which is due to core or monetary shocks but even that part of the *price level* due to what the authors call demand shocks.

This has an important consequence for the interpretation of Gartner and Wehinger's results. Core shocks in their study have permanent effects on the *price level* but not the inflation rate. Therefore, the link between the core inflation series constructed by the authors and inflation as a purely monetary phenomenon is much weaker. For theoretical reasons, a permanent change of the inflation rate must be caused by monetary factors, whereas a change of the price level may have various causes.

Apart form this conceptual problem, there arise some doubts about the validity of the identification scheme if one looks at the impulse response figures for Germany, the Netherlands and Italy. If the identification of the shocks is correct, one would expect that prices settle at a lower level after a positive supply shock. However, the price level in Germany returns to its original level, in the Netherlands does not react at all and in Italy the price level even rises. These responses indicate that supply shocks in those countries are systematically accommodated and that the model does not correctly distinguish supply and demand shocks.

In their extension of the Quah and Vahey approach to a trivariate VAR model including output growth, CPI inflation and short-term interest rate, the authors attempt a further breakdown of price changes into changes brought about by supply shocks, real demand and monetary shocks. As I said, I find this extension rather interesting because it might provide a direct way to identify that part of inflation, which is caused by monetary factors.

Clearly, this approach is interesting, because only that part of measured inflation which is caused by monetary factors can account for a sustained rise of the general price level and corresponds thus with our theoretical notion of inflation. In addition, from the viewpoint of a central bank the identification of inflation due to monetary shocks is important, since it would be the proper concept to judge the performance of a central bank.

Gartner and Wehinger identify monetary policy shocks by assuming that unexpected changes of the short-term interest rate are output neutral in the long run. At first sight, their empirical results support their interpretation of interest rate changes as discretionary policy instrument. For all countries, except for the Netherlands, the short-term interest rate is in the long run independent of all other variables in the system. A closer look, however, shows that the empirical findings are difficult to reconcile with economic theory. One would expect that a tightening of monetary policy or a rise of the short-term interest rate leads to a decrease of money demand, prices and possibly output. However, the results of Gartner and Wehinger show that in all countries except in the Netherlands, a permanently higher price level and a higher output follow a monetary contraction during the first year.

A final remark on the measurement of inflation concerns the price index used in the empirical analysis. Although Quah and Vahey and Gartner and Wehinger argue that the CPI is unsuitable for the measurement of inflation, both studies decompose inflation, measured by the CPI. The limited scope of the CPI and possibly its weighting may bias the measurement of underlying inflation. Clearly, if the method of Quah and Vahey identifies that part of measured inflation which is due to monetary shocks only and a monetary shock ultimately leads to a proportional rise in all prices, it should not matter which price index one uses for the decomposition. It would therefore be interesting to see how sensitive core inflation is with respect to the price index used in the exercise.

2. When is core inflation relevant for the monetary policy of the ECB?

Now I want to make three remarks on the question under what conditions the measurement of core inflation might be useful for the monetary policy of the ECB. First of all, in order to be useful, core inflation figures should be available with comparable frequency and speed as CPI inflation figures. This means that the VAR model should be estimated with monthly and not quarterly data.

Second, in order to be relevant at all in the short and medium term, the model should be estimated on aggregated European data. Indeed, Quah and Vahey and Gartner and Wehinger used a sample period of 25 years for their estimation. If one does not use aggregated European data, does the ECB has to wait 25 years for its first core inflation figure?

Finally, the core inflation series constructed by means of a VAR model might be sensitive to the sample period. Core inflation would be useless to monetary policy if every time new observations become available the core inflation series undergoes major revisions. Therefore, it should be investigated how sensitive the core series is to variations of the sample length.

Conclusions

My conclusions from this discussion of the study of Gartner and Wehinger are as follows: First, they should reconsider the empirical evidence on the stationarity of measured inflation. I suspect that the anomalies of the impulse response functions may be explained by the fact that core shocks have a permanent effect on inflation and not just the price level. As long as these anomalies persist it is not clear what component of measured inflation the identification scheme actually recovers. Second, one should experiment with price indices, defined for broader price sets and possibly without weighting. Moreover, in order to be useful for the ECB in the short and medium term, core inflation series have to be constructed using monthly, European data. Finally, I think it is important to find out how sensitive the core inflation series are to the length of the sample period. The concept will not become relevant to monetary policy if inflation figures undergo major revisions as soon as new observations become available.

Fiscal consolidation in general equilibrium models

Raf Wouters[1]

Introduction

In this paper, attention is focused on the effects of fiscal consolidation programmes in small open economies. The analysis concentrates on the intertemporal aspects of the problem: how will private sector behaviour be affected by public sector actions on public expenditure, taxes and deficits? Other aspects of fiscal programmes, such as the optimal tax decision, the possible distortions of taxes or the external effects of public capital formation, will not be discussed.

We start with a brief overview of the fiscal consolidation experience in Belgium. The magnitude and the specific content of the Belgian consolidation programme give it a special interest for those wishing to test some of the factors behind successful consolidation programmes. The specific characteristics of the Belgian experience become clear if they are compared with other European experiences.

In Section 2, we briefly review the literature on fiscal consolidation. Starting with a simple textbook model, different views on fiscal consolidation are discussed. The relative magnitude of the different channels of interaction between public and private sector behaviour will determine the success of the fiscal consolidation programme. In the literature, the experiences of Ireland and Denmark are considered as interesting examples of successful consolidation programmes. The simulation results from existing macroeconomic models are in most cases less optimistic, especially as far as the short-term effects of restrictive fiscal policies are concerned. General equilibrium models, although they rely on a totally different theoretical framework to traditional Keynesian models, also tend to yield negative output effects of fiscal consolidation programmes, at least insofar as the analysis is limited to the intertemporal aspect of the problem.

In Section 3, we present a theoretical general equilibrium model for a small open economy, so that the different aspects of the interdependence of public and private sector behaviour can be analysed in a coherent framework. This model allows us to analyse the sensitivity of certain effects to theoretical parameters: in particular, we can test the dependence of the result on the planning horizon and liquidity constraints of households, the importance of the labour supply reaction and the public consumption role in the utility function, the degree of price stickiness in an economy with imperfect competition, the exchange rate behaviour and the monetary policy reaction function.

In Section 4, we estimate a structural VAR model for Belgium, based on the theoretical insights of the general equilibrium model. The special contribution of fiscal shocks to economic growth and inflation (or the real exchange rate) will be estimated. The sensitivity of the results can be tested with alternative theoretical restrictions and empirical variables. Finally, we also apply the SVAR model to the Irish and Danish data, in order to identify the contribution of fiscal shocks to the growth process in these countries, and to compare the results with the Belgian experience.

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1 National Bank of Belgium, Research Department. The views expressed in this paper do not necessarily correspond to those of the NBB. The author wishes to thank P. Moës for his contribution to the empirical structural VAR analysis and M. Dombrecht for his comments on a previous version of this paper.

1. Belgian experience in fiscal consolidation

In Belgium, fiscal consolidation has been one of the central topics of economic policy since the beginning of the 1980s. The public deficit at the beginning of that decade was up to 12.8% of GDP, and public debt was rising quickly, from 60% in 1975 to more than 100% in 1982. Starting in 1982, there has been a gradual improvement in the situation of public finance, only temporarily interrupted at the beginning of the 1990s. The public deficit decreased and in 1997 Belgium is set to fulfil the Maastricht deficit criteria, with an estimated public sector borrowing requirement of 2.0% of GDP. Public debt is also declining, after reaching a peak of 135.2% in 1993; in 1997 it is estimated to be 122.2%.

This result was obtained despite a strong "snowball" effect. Interest charges increased from 6% of GDP in 1980 to 10.5% of GDP in the second half of the 1980s and the beginning of the 1990s. This implies that the improvement in the primary surplus shows an even stronger reversal: from -5% of GDP in 1981 to 5.9% of GDP in 1997 (see Figure 1).

These data are well known. What is less well known is that Belgium obtained these results almost exclusively via a decrease in government expenditure. Government revenue as a percentage of GDP has hovered around 46.7% over the period. Primary government expenditure, on the other hand, has decreased from 51.3% in 1981 to 41.7% in 1997, a level that is even below the European average. The effort that was made in terms of government expenditure cuts can be further illustrated by comparing the growth of real government expenditure in Belgium with the average European figures. In Belgium government expenditure, deflated by the consumer price index, increased by 15% over the period 1980-97, or 0.7% annually, against an average European growth of 2.3%.

All components of government expenditure have contributed to this result: government investment experienced the strongest decline over the period, with an average decrease of 4.8% in volume (European average +0.3%); purchases of goods and services decreased by 1.4% on average (+2.8% for the 15 EU countries); compensation of employees increased by 0.4% (1.4% for the 15 EU countries); subsidies to enterprises decreased by 0.8%, while the European average was +0.7%; and transfers to households increased by 1.9%, compared with an average of 3.1% for Europe.

During this long period of fiscal consolidation, economic growth in Belgium was slightly below the European average. However, the improvement in public deficits was accompanied by an amelioration in other macroeconomic fundamentals. Inflation and interest rates converged towards the German level, and the improvement in the public deficit also occurred simultaneously with the improvement in the current account balance.

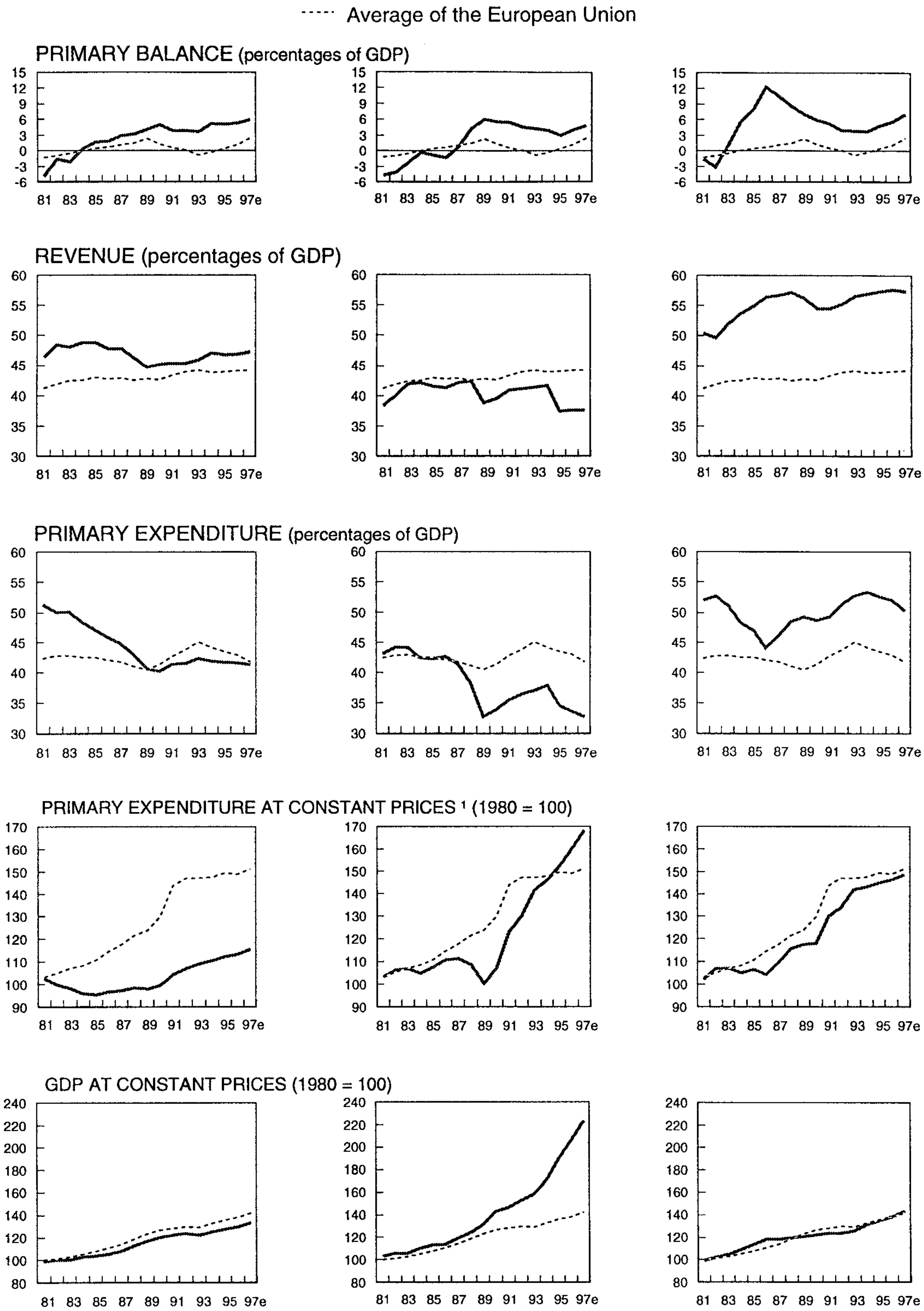
In the recent literature on fiscal adjustment, and especially the studies undertaken by the IMF (Alesina and Perotti (1996-97), McDermott and Wescott (1996)), Belgium is not cited as an example of successful fiscal consolidation, where the latter is defined as a period of tight fiscal stance such that the government debt/GDP ratio falls by at least 3 percentage points within two years, although, according to the most recent figures, Belgium qualifies for this definition over the period 1996-97. These studies have suggested not only that the size of the fiscal adjustment process is important (sharp contractions increase the probability of success) but also that the nature and composition of the measures are key elements for the success of the programme. Examples of such successful consolidation programmes were found in Ireland and Denmark, and these experiences are often discussed in the literature.

In Figure 1, we compare the Belgian experience with that of Ireland and Denmark. Although these countries experienced larger cuts in expenditure, over the whole period the consolidation effort was greater and more persistent in Belgium. In Section 4 these three examples of fiscal consolidation programmes are further analysed and, in particular, we try to estimate the specific contributions of public consumption cuts on observed economic growth and inflation.

Figure 1

Consolidation of public finances in some European countries

Belgium **Ireland** **Denmark**



[1] Deflated by the consumer price index.

Sources: EC, NAI and NBB.

2. A survey of the literature on the effects of fiscal consolidation

In the literature one finds many simulation exercises for fiscal policy shocks using existing econometric and more theoretically oriented models. As fiscal consolidations affect in the first instance the intertemporal constraints on private sector behaviour, we focus especially on those studies that incorporate forward-looking expectations for the expected income streams, as only these models can take into account the effects of announced fiscal policies.

Fiscal policy affects economic growth via many different channels. In this section we review those channels that have received most attention in the literature. Starting with more theoretical arguments on planning horizon, expectations and labour supply reaction, we move on to the more practical or empirical questions such as the rigidity of prices, exchange rate behaviour and monetary policy reactions. While the second category of topics is crucial in explaining the short-run impact of fiscal adjustments, the theoretical arguments remain important as they determine the long-run reaction of models to fiscal shocks. In the next sections some of these topics will be further discussed within the framework of a small general equilibrium model for an open economy.

2.1 The time horizon of the private sector

In the traditional Keynesian models, consumption is determined by current income. The outcome of a fiscal consolidation in this case is simple. A decrease in fiscal spending has contractionary real effects in terms of both consumption and output, because it shifts the aggregate demand curve downwards and prices are rigid. Consumers do not perceive the positive effects of lower public deficits on future taxes and therefore on wealth. The typical multiplier analysis implies that the effect of government spending cuts will be larger than the effect of tax increases.

In the neoclassical Ramsey model, agents have an infinite time horizon and the labour supply is inelastic. In this model, consumption is proportional to wealth, defined as the present value of all future revenue. Lower taxes and debt financing will leave wealth unaffected as consumers discount the higher expected future taxes. A permanent change in public consumption financed with lump-sum taxes results in an equal but opposite change in private consumption, as consumption moves proportionally with wealth. So aggregate demand, output and employment are unaffected. As consumers are forward-looking, they recognise the wealth effect resulting from the change in the present value of future taxes necessary to finance public expenditure. This neutral effect of fiscal spending on economic activity is a result of the Ricardian equivalence hypothesis (Barro (1974)).

The hypothesis of infinite horizons is however a strong hypothesis. In reality, consumers have finite lives and they typically save for consumption in later periods of the life cycle. So unless intergenerational transfers are given a similar valuation to consumption in the utility function, household horizons will be less than infinite. In modern macro-modelling, this finite horizon assumption is typically introduced by using a constant probability of death or the perpetual youth hypothesis (Blanchard (1985)). This approach implies that future taxes are discounted at a higher discount rate (the discount rate plus the expected probability of dying), and therefore receive a lower expected value, as the consumers are uncertain whether they will still be alive at the time the future taxes are levied. Under this hypothesis, government taxes will have an impact on wealth, and therefore on private consumption. For the same reason, public spending cuts will be less than fully offset by the increase in private consumption and declining aggregate demand will cause lower real interest rates. This will stimulate investment and increase the optimal capital stock. Given the hypothesis of fixed labour supply, output will increase with permanent public spending cuts.

Finite horizons are not the only reason for an absence of Ricardian equivalence. Other reasons are imperfect capital markets and liquidity constraints. Empirical consumption functions show in most cases an excessive sensitivity with respect to current income. Therefore, an alternative (or complementary) solution to make economic models more realistic is to define part of the households as being liquidity constrained. Such market imperfections enhance the real effects of fiscal policy, as we will see in our model discussion later.

Up to now, we have considered only permanent changes. Transitory changes have a smaller effect on wealth but they entail an additional intertemporal substitution effect as they may affect the rate of interest. A temporary spending cut will have a smaller impact on wealth compared with a permanent one, increasing private consumption less than proportionally, and so lowering aggregate demand and, therefore, the interest rate. Lower interest rates will shift consumption towards the present period and at the same time stimulate capital accumulation. So a transitory contractionary fiscal programme has a positive effect on output as the capital stock increases. As public spending subsequently returns to its original level, private consumption will decline less, causing a temporary increase in interest rates and lower investment, so that capital stock, output and consumption will return to their original levels.

The main conclusion here is that the introduction of the intertemporal budget constraint for households, in contrast to the current income approach in Keynesian models, lowers the multiplier effects of permanent government spending or tax changes. But two further remarks should be made.

Although these models are able to yield expansionary fiscal contractions, the logic behind the results contradicts recent ideas behind expansionary fiscal consolidation programmes. It is the decline in aggregate consumption that lowers the interest rates and stimulates capital accumulation and output gradually. Transitory spending shocks, which have larger impacts on interest rates, will be more effective in the short run. The modern view on expansionary fiscal consolidation, on the contrary, stresses the importance of the permanent character of fiscal programmes to generate positive effects on expectations, private consumption and investment demand. To achieve such expansionary expectations effects, one has to introduce specific assumptions as to expectation formation that rely on uncertainty about the future action of the public sector and on non-linear reaction functions of the public sector to unsustainable fiscal programmes. Futhermore, the results obtained in the present section were derived under the hypothesis of constant employment equal to the fixed labour supply. These two remarks will be further discussed later on.

2.2 The impact on expectations: credibility, persistence and composition of the fiscal programme

With regard to the deep recession in West Germany in 1981-82, with historically high public deficits (4.9% in 1981 and 4.4% in 1982), followed by a quick recovery during a period of restrictive fiscal policy in 1983-86, some German economists have put forward the hypothesis that economic growth was strongly influenced by the expectation effects of fiscal policy (Fels and Froehlich (1986), Hellwig and Neumann (1987)). Rapidly growing deficits may undermine private sector confidence in the future economic outlook, causing a decline in consumption and investment and leading to higher interest rates. Fiscal austerity, on the other hand, could stimulate a "psychological crowding-in", given public approval of a policy aimed at long-term stability. By absorbing a smaller share of GDP, the public sector made room for the private sector to expand. In particular, more savings could be channelled into productive private investments. The decrease in the government borrowing requirement, together with a strict anti-inflationary monetary policy, paved the way for a substantial decline in interest rates.

The Danish (1982-86) and Irish (1987-89) experiences gave further support to this hypothesis. Giavazzi and Pagano (1990) explained the strong recovery in both countries after their fiscal adjustment programmes by the "German view": a fiscal shock can trigger a positive reaction of private demand, lower real interest rates and create positive wealth effects. Giavazzi and Pagano (1995) later applied their explanation to a broader set of OECD countries. The same view underlies the work by Alesina and Perotti (1996-97), which seeks to define a "successful fiscal consolidation plan".

Given these observations, some theoretical justifications for stronger-than-proportional wealth effects following a fiscal adjustment programme were developed. Blanchard (1990) argues that a tax increase today can have expansionary effects if it generates expectations of less dramatic and disruptive tax increases tomorrow. By removing the uncertainty about the evolution of future fiscal policy, it may reduce precautionary saving.

Bertola and Drazen (1993) introduce trigger points for the government expenditure/output ratio at which sharp policy changes are expected to occur. Such trigger points make government expenditure follow a discrete process. Once expenditure reaches the critical level, a drastic stabilisation programme is expected. So if public spending comes close to a trigger level, a further rise in expenditure will increase the possibility of a drastic future stabilisation programme, and will therefore lead to a decrease in the expected future expenditure stream. As a critical level is reached, either the stabilisation programme will be put into action and government spending will be cut while consumption will increase, or the expected stabilisation process will not be implemented, and then consumption will make a negative jump and a higher trigger level for public spending will be established. Such a process can explain the Danish and Irish cases to the extent that the stabilisation programmes were effectively expected by the household sector.

These models introduce non-linear effects based on the change in the perception of uncertain future fiscal policies. Increases in public debt in a context of an already high level of debt can have different effects to an increase of debt at a low level. At a high debt level, a further increase makes a shock programme with drastic measures more likely because the continuation of past policies becomes unsustainable. In such a case, a further deficit expansion not only raises future taxes proportionally but, as it also increases the probability of a drastic programme in the near future with a finite time horizon (as in the model of Sutherland (1995)), this also leads to a more-than-proportional fall in future income. So a fiscal expansion can have contractionary effects on aggregate demand, while a fiscal contraction can lead to an expansion in aggregate demand.

Another possible channel that can increase the strength of the wealth effect is the decrease in interest rates that follows the stabilisation process. Interest rates can fall if monetary policy follows a rather expansionary stance to compensate for the restrictive fiscal policy. This effect will be addressed in a later section, where the interaction between fiscal and monetary policy is further discussed. Long interest rates can also decrease because the required risk premium declines. Falling public debt will diminish the danger of its monetisation and, therefore, the inflation bias due to public debt devaluation. The restored confidence can also attract foreign investors looking for interesting capital gains on high-yield bonds. A decrease in public debt will also reduce the default risk as it minimises the probability of an unsustainable snowball effect.

The lower interest rates will increase the value of financial assets, equity and house prices. Together with the optimistic expectations about future growth prospects in the private sector, this can lead to a boom in investment demand in both housing and business sectors.

The strength of the wealth effect depends strongly on the private sector's perception of future fiscal policy. Some authors therefore stress that the size and specific composition of the fiscal programme can be important in bringing about the necessary confidence shift (Alesina and Perotti (1996-97)). The size of the fiscal measures taken is important as it gives a signal on the unobservable future course of the process. A small adjustment programme can disappoint the private sector and will therefore not cause the necessary jump in expectations and private expenditure. Following this logic, it is also important to demonstrate that the measures taken are permanent measures, to signal clearly that a significant break has occurred with the past process of public deficits and expenditure. The composition of the adjustment programme can be important in this context. As some expenditure cuts, for instance transfers and compensation cuts, are politically more difficult to implement, they will be more convincing in signalling the willingness for further changes. Alesina and Perotti and McDermott and Wescott provide evidence which shows that successful stabilisation programmes are typically of larger magnitude and share a similar composition, with a preference for spending cuts and an aversion to income tax increases. The authors cite the Irish experience of 1987-89 and the Danish programme of 1982-86 as typical examples of successful programmes.

The empirical relevance of this "expansionary fiscal contraction" is difficult to prove, as the argument is strongly based on the behaviour of unobservable expectations. The non-linear nature of the proposed relation also makes it difficult to model the hypothesis and to test it empirically. One approach for identifying the presence of these expectation effects is to estimate consumption and investment functions and check whether the observed reaction in private spending corresponds to the normal behaviour of consumers (Giavazzi and Pagano (1990, 1996)). Unexplained positive shocks in private demand around periods of fiscal contraction are then considered as evidence of positive expectation effects. This approach can give an indication of a structural break in the consumption function, but it cannot prove the hypothesis or the specific channel of the expectation effect. Barry and Devereux (1995) show that alternative interpretations for the remarkable results of Ireland and Denmark are possible. These can be found in the presence of other shocks affecting the economy at the time the fiscal shock took place. Indeed, the fiscal consolidation programme was preceded in both countries by a real depreciation that improved competitiveness and net exports. The fiscal measures were also accompanied by a shift in monetary policy. Both countries shifted towards a policy of a stronger currency, which may have influenced expected inflation and interest rates. Furthermore, real wages and labour markets experienced shocks at the same time as the fiscal programme was being carried out. So there is a clear problem of disentangling different shocks that were occurring simultaneously, and this makes interpretation of these experiences very difficult. A more general and structural approach is necessary to identify alternative shocks and interpret the joint observation of different macroeconomic variables. In our SVAR experiments, we try to take a step in that direction.

Most empirical macro models simulate the deflationary effects of government cuts on aggregate demand, although some arrive at rather small-short term costs (for instance the IMF Multimod exercise on Canada by Bayoumi and Laxton (1994)). But such results are mainly due to other channels. As these economies are small and open, competitiveness should play a dominant role in the interpretation of the simulation results. At the same time, the monetary policy reaction and its effects on the exchange rate are crucial for the outcome of such measures. These points will be taken up in later sections.

Up to now, the discussion has focused on demand effects. In the next two sections we discuss the effects on the supply side of the economy: household labour supply and firm's demand for labour.

2.3 The effect on the labour supply

In Section 2.1, following the classical Ramsey model, the labour supply was assumed to be inelastic. In general equilibrium models, the utility of households depends on both consumption and leisure. The labour supply then reacts to changes in consumption or wealth. In Barro (1989), Baxter and King (1993) and Aiyagari, Christiano and Eichenbaum (1992), the increase in private consumption and wealth, following public spending cuts, also stimulates a higher demand for leisure and therefore gives rise to a negative labour supply effect. This results in higher private consumption but lower employment and output in the new steady-state equilibrium. As the equilibrium capital stock and investment also decrease, the total multiplier effect in these general equilibrium models can easily exceed one. A permanent reduction has greater output effects than a temporary one, because the wealth effects, and therefore the impact on the labour supply, are larger with permanent measures. These results are remarkable as they reproduce the Keynesian result in a neoclassical framework of full employment and flexible prices.

In practice, such wealth effects on the labour supply may be rather limited. Furthermore, negative labour supply effects of fiscal consolidation may be compensated for by a positive impact of lower distortion effects of taxes on employment and investment as, in reality, taxes are not of a lump-sum type. Income taxes in neoclassical models (Baxter and King (1993)) cause a divergence between real wage costs for firms and the disposable wage for employees. By driving a wedge between private and social returns on labour, employment will be lower in an economy with labour tax distortions than in the optimal world with lump-sum taxes. Fiscal consolidation in these models will therefore have a positive supply effect by eliminating tax distortions.

These different steady-state supply effects of both assumptions are also present in the simulation results of modern macroeconomic models that are constructed around a well-defined steady-state model. The simulation results of the Quest II model (Roeger and In't Veld (1997)), for instance, show how the impact of a fiscal consolidation programme depends on the financing decision. If the decrease in public spending goes together with a reduction in lump-sum transfers, the long-term effect on output is negative. If the spending cuts give room for a decrease in labour taxes, there will be a significant positive effect on employment. The importance of the financing decision was also stressed by Bartolini, Razin and Symanski (1995), using the IMF Multimod model to simulate fiscal restructuring in the G-7 countries.

The relative size of these different effects in empirical studies is ambiguous. Studies on the labour supply give different results according to whether they are based on micro- or macroeconomic data. The size of distortion effects will, in addition, depend on the structure of the labour market. Labour tax distortions can also be offset by the external effects of public expenditure and investment on private labour and capital productivity. In this paper, however, we concentrate on the intertemporal aspects of fiscal consolidation, which should be distinguished from the other aspects of fiscal policy, no matter how important the structure of the financing decisions and of the expenditure composition may be.

The effects of fiscal contractions on the labour supply may also be different in models with labour unions and bargaining. Changes in the system of accommodating transfers, especially those related to unemployment, can change the insider behaviour of unions in the labour market (Calmfors and Horn (1985-86)). Abolition of the automatic indexation of transfers and public sector compensation can have spillover effects on the private sector wage formation process. These channels are interesting as they offer a different view on expansionary fiscal consolidations. They suggest that the negative demand effects of spending cuts may not be offset by private demand shocks, but rather that positive supply shocks could be important. As supply shocks tend to have more permanent effects on output, this distinction can have major consequences for the long-run outcome as well. In the SVAR exercise, we will be able to discuss the offsetting role of these different channels.

2.4 Imperfect competition models with sticky prices

Up to now, the models discussed have all assumed flexible prices and full employment. Under these neoclassical assumptions, demand shocks do not affect the supply decisions of the firms directly, and the short-term effects of public finance shocks on output are thus rather small. In reality, however, the decision to cut government expenditure is often taken in a context of unemployment, and spending cuts are often postponed because of the expected short-run costs in terms of negative employment effects in a situation of already high unemployment.

Therefore, the problem of fiscal consolidation should be examined using a model with unemployment and rigid prices or wages, so that demand shocks do have short-run output effects.

Perfectly competitive models predict that aggregate demand shocks can raise output and employment only by increasing households' willingness to supply labour (Woodford and Rotemberg (1992)). A profit-maximising firm in perfectly competitive output and factor markets produces and hires labour until marginal productivity equals marginal costs. As these variables depend only on supply-side conditions in the form of installed capital, technology, etc., demand shocks will not affect the real output and employment decisions of the firm. The demand for labour will shift only if the assumption of perfect competition and price flexibility is dropped. The neoclassical models, by assuming perfect competition and flexible prices, do not allow for the effect of fiscal spending shocks on the supply decisions of the firms and labour demand in particular. In neglecting these effects, the neoclassical model underestimates the short-run costs of fiscal adjustments.

Imperfectly flexible wages can explain the influence of aggregate demand shocks on real output, but they imply countercyclical real wages. Therefore, it is more realistic to look for a solution in terms of imperfect competition in the goods market as the rationalisation for the impact of demand shocks on firms' supply decisions. Monopolistic competition in the goods market also implies equilibrium situations with unemployment. Price rigidity can be rationalised in these environments both as a consequence of collusive behaviour between oligopolies (Rotemberg and Woodford (1992)) or in terms of price adjustment costs (Calvo (1983), Kollman (1997) and Hairault and Portier (1993)).

In such models, aggregate demand shocks affect the mark-up as prices do not fully reflect the increase of marginal costs. The reaction of output will no longer depend exclusively on the labour supply reaction of households; labour demand by firms will also shift. These neo-Keynesian models succeed in combining price rigidity and the importance of the demand shocks in the short run within the long-run neoclassical framework.

Empirical macroeconomic models typically incorporate sticky price and wage assumptions in the short run, and are therefore demand-driven in the short run. They are able to illustrate important short-run output costs of public spending cuts. The problem, however, is that such models, by incorporating more realistic, empirically estimated short-run dynamics, lose their theoretical consistency, especially in terms of the profit maximisation behaviour of firms in output and price decisions. This was certainly the case in traditional macroeconomic models, inspired by the old Keynesian view and lacking the long-run steady-state framework that is needed to determine long-run stock flow equilibrium. But even more modern models, built around a theoretical steady-state model, have a somewhat arbitrary combination of short-run dynamics and long-run steady-state properties.

2.5 Open economies and monetary policy reaction functions

The simple Mundell-Flemming approach implies that fiscal spending cuts will decrease aggregate demand and result in a real depreciation of the exchange rate together with an improvement in the current account.

This result remains more or less valid in the modern approach using rational expectations and intertemporal optimisation. In a small open economy model with one good and finite horizons, aggregate demand declines with spending cuts. As the interest rate is fixed for the small economy, contrary to the closed economy case, investment does not react but net exports increase. The accumulation of net foreign assets substitutes for the accumulation of domestic capital, and therefore output remains unchanged. In the ultimate steady state these effects will be reversed, allowing a further increase in private consumption, as equilibrium in the balance of payments requires the trade deficit to compensate for the increase in interest income on net foreign assets.

In a model with two goods and imperfect substitution, there will be additional effects via changes in the real exchange rate, as the price of the domestic good in terms of the foreign good will decrease. The real depreciation goes hand in hand with a decrease in the real interest rate so that investment and output rise, a result that compares well with the one obtained in the closed economy case. Further complications of the model (e.g. Ahmed (1987) and Cuddington and Vinals (1986)), introducing tradable and non-tradable goods and wage-price inflexibility, can, however, make these results ambiguous. The non-tradable sector output declines with the public spending cuts, and real wages go down, boosting exports in the tradable sector.

In practice, the behaviour of the nominal exchange rate depends strongly on the reaction of monetary policy and the change in risk premia for small open economies.

Bayoumi and Laxton (1994), using simulations of the IMF Multimod model for Canada, illustrate how the outcome of a fiscal consolidation programme depends on the interaction between the exchange rate and monetary policy. A deficit reduction package that uses a combination of increases in taxes less transfers and a decrease in government expenditure to bring the debt/GDP ratio down was simulated for Canada. Monetary policy responds endogenously as it pursues an inflation target and long-term interest rates incorporate a small risk premium that depends on the public debt ratio. In this simulation, the size of the short-run costs depends on the perception by economic agents of the persistence of the deficit cut. With a fully credible fiscal programme, economic agents anticipate the decline in future interest rates, and the exchange rate depreciates immediately. Exports rise and the total output effect can even become positive. If the programme is not credible, the exchange rate depreciates less as economic agents do not correctly anticipate future lower interest rates. Therefore, output and inflation decline initially. This example illustrates the importance of the currency depreciation and a monetary policy rule in determining the outcome of consolidation programmes for small open economies. An increase in competitiveness and net exports can be an important channel to offset negative domestic demand shocks in the short run.

This result is typical for many simulations on fiscal consolidation. In a discussion of the impact of the Maastricht criteria and the deficit reduction programmes on economic growth, Buiter (1993) also points to the dependence of the short-run costs on the behaviour of interest rates and exchange rates to offset the negative impact on demand. In that context, the disappearance of long-term interest rate differentials with Germany did play an important role in softening the short-run costs.

In addition, in the QUEST model (Roeger and In't Veld (1997)) the short-run effects of fiscal contractions depend strongly on the monetary policy rule. With a nominal interest rate target, money supply decreases following spending cuts and this will enforce the negative short-run costs. On the other hand, if monetary policy follows a strict money supply rule, nominal interest rates decline and such a type of monetary policy can even reverse the short-run effects.

3. Fiscal consolidation in a general equilibrium model for a small open economy

In this section, a small general equilibrium model for an open economy is presented. The model integrates most of the topics that were discussed in the previous section, and therefore allows us to analyse the importance of such effects in a coherent framework. After a brief description of the model, the impulse response effects of public spending cuts and tax increases are presented and the sensitivity of the results with respect to the parameter values is tested.

3.1 The household sector

A first group of households is liquidity constrained and has no access to the capital market. These households consume disposable labour income during the period in which it is earned, and it is supposed that their labour supply is perfectly elastic so that it fluctuates together with total employment, determined elsewhere in the model.

The behaviour of these households is summarised by the following equation:

$$C_t^w = \frac{w_t}{p_t}(1-L)_t - T_t \tag{1}$$

where income is equal to labour income minus net taxes. Notice that taxes are treated as lump-sum taxes and are therefore not proportional to income. The question of tax distortions is not considered in this model.

The second group of households has full access to the capital market: they hold money balances (M), domestic government bonds (B), foreign interest-bearing bonds (F) and domestic equity

(V). Bonds are one-period assets on a discount basis, such that the price in period t of the domestic bond (b) equals $1/(1+R)$ and, similarly, for the foreign bond with price (f):

$$\frac{M_{t+1}}{p_t}+\frac{b_t B_{t+1}}{p_t}+s_t\frac{f_t F_{t+1}}{p_t}+\frac{d_t V_{t+1}}{p_t}= \tag{2}$$
$$(1/pr)\left[\frac{M_t}{p_t}(1+gm_t)+\frac{B_t}{p_t}+s_t\frac{F_t}{p_t}+\frac{d_t V_t}{p_t}+\frac{w_t}{p_t}(1-L)_t-C_t^1-C_t^2-T_t\right]$$

pr is the probability of survival. A perfect insurance market inherits consumers' wealth on their death and redistributes wealth in the form of an annuity payment in proportion to household wealth. V stands for domestic equity with price d. Real wealth (W) is equal to:

$$W_t=\frac{M_t}{p_t}(1+gm_t)+\frac{B_t}{p_t}+s_t\frac{F_t}{p_t}+\frac{d_t V_t}{p_t} \tag{3}$$

The utility function is of the following type:

$$V_t=a\ln C_t^1+(1-a)\ln\left[C_t^2+\theta G_t\right]+\frac{1}{1-\varpi}L_t^{1-\varpi} \tag{4}$$

Utility depends on the consumption of cash ($C1$) and credit goods ($C2$), public consumption (G) and leisure time (L).

The households have a discount factor (β) and a finite expected life, with pr the probability of survival and a lifetime horizon $1/pr$, so that the objective function becomes:

$$\max \underset{t}{E}\sum_j(\beta pr)^j V\left[C_{t+j}^1,C_{t+j}^2,G_{t+j},L_{t+j}\right] \tag{5}$$

and the cash constraint applies to the consumption of cash goods:

$$C1_t\leq\frac{M_t}{p_t}(1+gm_t) \tag{6}$$

The first-order conditions are derived from the Lagrangian, combining the optimisation function and the constraints, with λ the Lagrange parameter for the budget constraint and η for the cash constraint:

$$V_t^{C1}=\lambda_t+\eta_t \tag{7}$$

$$V_t^{C2}=\lambda_t \tag{8}$$

$$V_t^1=\frac{w_t}{p_t}\lambda_t \tag{9}$$

$$E_t\left[\frac{\beta pr}{\gamma}\frac{\eta_{t+1}}{p_{t+1}}+\frac{\beta pr}{\gamma}\frac{\lambda_{t+1}}{p_{t+1}}\right]=\frac{\lambda_t}{p_t} \tag{10}$$

$$E_t\left[\frac{\beta pr}{\gamma}\frac{\lambda_{t+1}}{p_{t+1}}(1+R_t)\right]=\frac{\lambda_t}{p_t} \tag{11}$$

$$E_t\left[\frac{\beta pr}{\gamma}\frac{\lambda_{t+1}}{p_{t+1}}\frac{s_{t+1}}{s_t}(1+R_t^f)\right]=\frac{\lambda_t}{p_t} \tag{12}$$

$$E_t\left[\frac{\beta pr}{\gamma}\frac{\lambda_{t+1}}{p_{t+1}}\frac{d_{t+1}}{d_t}\right]=\frac{\lambda_t}{p_t} \tag{13}$$

γ stands for technological progress (but it is assumed to equal 1 in the rest of the analysis).

Equations (7) and (8), combined with the first-order constraint on cash holdings (10) and the interest rate condition (11), result in a velocity of money that depends positively on the interest rate. Using the equality between $C1$ and real money holdings (M/P), the equation can then be rewritten as:

$$C_t \Big/ \frac{M_t}{p_t}(1+gm_t) = \frac{1}{a} + \frac{1-a}{a} R_{t-1} \tag{14}$$

In the simulation of the model, this money demand equation is specified in terms of total consumption and not just in terms of the consumption of the unconstrained consumers.

Equations (11) and (12) represent the uncovered interest rate parity condition for nominal exchange rate determination. Equation (13) shows that the expected holding return on equity equals the expected one-period interest rate under certainty equivalence.

The consumption of the second type of household can be approximated by using the first-order conditions and the restriction that the net present value of consumption must equal total expected revenue and actual wealth. The consumption of cash goods becomes:

$$C1_t = (1-\beta pr/\gamma)a\frac{1}{(1+R_{t-1})}\left[W_t + R_{t-1}\frac{M_t}{p_t}(1+gm_t) + H_t + \theta PG_t\right] \tag{15}$$

and the consumption of credit goods, including government consumption, is related to the consumption of cash goods in the following way:

$$C2_t^T = C2_t + \theta G_t = \left(\frac{1-a}{a}\right)(1+R_{t-1})C1_t \tag{16}$$

Total aggregate consumption can then be expressed as a function of total wealth:

$$C_t^T = \beta_t^0\left[W_t + R_{t-1}\frac{M_t}{p_t}(1+gm_t) + H_t + \theta PG_t\right] \tag{17}$$

where $\beta_t^0 = (1+\frac{1-a}{a}(1+R_{t-1}))\frac{(1-\beta pr/\gamma)a}{1-(1-\beta pr/\gamma)a}\frac{1}{(1+R_{t-1})}$ and H and PG stand for the present value of disposable labour income and government expenditure discounted with $pr/(1+RR)$.

Substituting out human wealth and the present value of government expenditure, the equation can be rewritten as:

$$C_t^T = \frac{(1+RR_t)(\beta_t^0/\beta_{t-1}^0)\beta/\gamma}{(1-(1-\beta pr/\gamma)a)}C_{t-1}^T - \frac{(1-pr)}{pr}\beta_t^0\left[W_t + R_{t-1}\frac{M_t}{p_t}(1+gm_t)\right] \tag{18}$$

The introduction of two types of household and a finite horizon allows a generalised permanent income approach. The liquidity constraint on part of the consumers can explain the excess sensitivity of consumption to current income innovations. The finite time horizon assumption allows testing of the impact of different planning period hypotheses on the effect of public deficits on private consumption. By incorporating government consumption and private consumption in the consumers' utility function, the results of different assumptions about the substitution or complementarity between both types of goods can be analysed. A negative value for θ implies that an increase in government consumption raises the marginal utility of private consumption (i.e. the two are complements), whereas a positive θ suggests that an increase in government consumption diminishes the marginal utility of private consumption (i.e. the two are substitutes).

In the case of monopolistic competition, aggregate consumption has to be considered as an index of many different consumer goods. The allocation is considered in two steps together with the other final demand components: investment and government consumption. First, aggregate demand is allocated between domestic and foreign goods and then between an infinite series of differentiated domestic or foreign goods. The final demand index is defined as a CES function $(C+G+I) = (H^{1/(1+sa)}+F^{1/(1+sa)})^{(1+sa)}$, where H is an index of consumption goods produced in the country, and F is an index of imported goods. The final demand price index is defined as: $P = (1/(pH^{-1/sa}+pF^{-1/sa})^{(-sa)}$. The optimal share is $YH/(C+G+I) = (pH/P)^{-(1+sa)/sa}$.

H is an index of domestic goods h. There exists a continuum of home-produced goods, indexed by s element of [0,1]. So yH, the demand for home goods by domestic (H) and foreign sources (exports), and yF, the total demand of import goods, can be defined as follows:

$$y_t^H = \left[\int_0^1 h_{j,t}^{1/(1+\upsilon)} dj\right]^{1+\upsilon} \tag{19}$$

$$y_t^F = \left[\int_0^1 f_{j,t}^{1/(1+\upsilon)} dj\right]^{1+\upsilon} \tag{20}$$

with $1+\upsilon/\ \upsilon$ denoting the price elasticity of demand. The optimal consumption allocation implies:

$$y_{j,t}^H = (p_{j,t}^H / p_t^H)^{-(1+\upsilon)/\upsilon}(y_t^H / n) \tag{21}$$

$$y_{j,t}^F = (p_{j,t}^F / p_t^F)^{-(1+\upsilon)/\upsilon}(y_t^F / n) \tag{22}$$

3.2 The firm problem

We assume that firms have some market power and behave as monopolistic competitors. The model allows increasing returns to scale (either in the form of overhead costs or in terms of externality). Firms use labour, capital and energy inputs. Energy is used in a fixed proportion to output.

Firm j maximises its expected profit, discounted with a rate (ρ) which is determined by the valuation of the shareholders (the unconstrained consumers). ρ_{t+1} can be replaced by the shadow value of wealth λ_{t+1} of households.

The model of price determination, inspired by Calvo (1983), assumes that firms are not allowed to change their prices, unless they receive a random "price change signal". The probability that a given price can be changed in any particular period is constant (1-ς). This probability also determines the fraction of all prices that are changed in each period. Consider now the problem for firm j, which is allowed at time t to set a new price p_j. At the time that firm j changes its price, there are three control variables p_j, H_j and K_j. Firm j will maximise the following expectation:

$$E\sum_t\sum_i \beta^{t+i}\rho_{t,t+i}\left[\begin{array}{l}\zeta^i p_j^H y_{j,t+i} + (1-\zeta^i)p_{j,n}^H y_{j,t+i} - w_{t+i}H_{j,t+i} - p_{t+i}^e s_{t+i}\dfrac{i_e}{1+i_e}y_{j,t+i} \\ -p_{t+i}K_{j,t+1+i} + p_{t+i}(1-\tau)K_{j,t+i} - p_{t+i}\dfrac{\psi}{2}\dfrac{(K_{j,t+1+i}-K_{j,t+i})^2}{K_{j,t+i}}\end{array}\right] \tag{23}$$

subject to the production technology and the demand for good j:

$$y_{j,t+i} \le (1+i_e)A_{t+i}K_{j,t+i}^{\alpha\mu}H_{j,t+i}^{(1-\alpha)\mu}e^{\gamma t} \tag{24}$$

$$y_{j,t+i} \le y_{j,t+i}^H = \left(\frac{p_{t+i}^H}{p_j^H}\right)^{\frac{1+\upsilon}{\upsilon}}\frac{y_{t+i}^H}{n} \tag{25}$$

The first-order condition for labour is:

$$w_t = (1+i_e)F^H_{j,t}(p^H_{j,t} - v_{j,t}) - p^e_t s_t i_e F^H_{j,t} \tag{26}$$

or, introducing the real marginal cost:

$$\frac{w_t}{(1-i_e)p^H_{j,t}F^H_{j,t}} + \frac{p^e_t s_t i_e}{(1-i_e)p^H_{j,t}} = mc_{j,t} \tag{26'}$$

The first-order condition for capital is:

$$\rho_t p_t (1+\psi\frac{(K_{j,t+1+i} - K_{j,t+i})}{K_{j,t+i}}) = \beta\rho_{t+1}p_{t+1} \left[1-\tau+(1+i_e)F^K_{jt}(1-\frac{v_{j,t+1}}{p^H_{j,t+1}}) - \frac{p^e_{t+1}s_{t+1}i_e}{p^H_{j,t+1}}F^K_{jt} + \psi\frac{(K_{j,t+2+i} - K_{j,t+1+i})}{K_{j,t+1+i}}\right] \tag{27}$$

By introducing Tobin's Q into the demand for capital, equation (28) can be simplified:

$$\rho_t p_t Q_t = \beta\rho_{t+1}p_{t+1}\left[Q_{t+1} - \tau + (1+i_e)F^K_{jt}(1-\frac{v_{j,t+1}}{p^H_{j,t+1}}) - \frac{p^e_{t+1}s_{t+1}i_e}{p^H_{j,t+1}}F^K_{jt} + \psi\right] \tag{27'}$$

And the condition for the price is:

$$\sum_i \beta^{t+i}\rho_{t+i}\varsigma^i\left(\frac{p^H_{t+i}}{p^H_j}\right)^{\frac{1+\upsilon}{\upsilon}}\frac{y^H_{t+i}}{n} = \sum_i \beta^i\rho_{t+i}\varsigma^i v_{j,t+i}\frac{1+\upsilon}{\upsilon}\left(\frac{p^H_{t+i}}{p^H_j}\right)^{\frac{1+\upsilon}{\upsilon}-1}\left(\frac{1}{p^H_j}\right)p^H_{t+i}\frac{y^H_{t+i}}{n} \tag{28}$$

Using (26) and (27), the price at time t can be derived as:

$$p^H_j = \frac{\sum_i \beta^{t+i}\rho_{t+i}\varsigma^i(1+\upsilon)(p^H_{t+i})^{\frac{1+\upsilon}{\upsilon}}\frac{y^H_{t+i}}{n}\left[\frac{w_{t+i}}{(1+i_e)F^H_{t+i}} + p^e_t s_t i_e\right]}{\sum_i \beta^{t+i}\rho_{t+i}\varsigma^i(p^H_{t+i})^{\frac{1+\upsilon}{\upsilon}}\frac{y^H_{t+i}}{n}} \tag{28'}$$

which shows clearly that the price set by firm j, at time t, is a function of expected future marginal costs. The price will be a mark-up over marginal costs. If prices are perfectly flexible, the mark-up will be constant and equal to $(1+\upsilon)$. With sticky prices, the mark-up becomes variable over time when the economy is hit by exogenous shocks. A cost increase temporarily lowers the mark-up such that production is less affected than in the flexible price case. A positive demand shock also lowers the mark-up and stimulates employment, investment and real output. Through this last channel the model obtains a Keynesian character: following a government demand shock, firms are stimulated to increase production. This contrasts with the classical real business cycle tradition, where the supply reaction of firms is not directly affected by demand shocks. The introduction of increasing returns to scale can further enhance the supply reaction of firms following a demand shock.

3.3 The government sector

The government sector has to satisfy the following budget restriction:

$$\frac{b_t B_{t+1}}{p_t} = \frac{B_t}{p_t} + G_t - T_t \tag{29}$$

The primary deficit (G-T), together with the debt servicing has to be financed by the issuance of new public debt B at price b. To prevent the public deficit becoming explosive, the following endogenous tax behaviour is assumed:

$$T_t = g(\frac{B_t}{p_t} - \frac{B_t^0}{p_t}) + \varepsilon_t^T \tag{30}$$

with g greater than the real interest rate minus real growth ($g > R$-π-γ), so that stable public debt is guaranteed at the long-term objective B^0. ε^T represents stochastic tax shocks. The effect of the tax shocks will depend on the specification of the time horizon and the liquidity constraints on consumers. In the most simple specification with infinite horizon and no liquidity constraints, taxes will be a perfect substitute for debt financing and the size of parameter g will have no impact on the dynamics of the model. With finite horizons and liquidity constraints, the impact of taxes becomes more complicated since it influences the households' budget constraints.

Government expenditure affects the budget constraint on the private sector via the wealth effect, even with infinite household time horizons. But in that case the financing decision becomes irrelevant. With finite horizons or liquidity restrictions, the financing decision will make a difference.

3.4 The balance of payments and foreign demand

The accumulation of foreign assets (F) is determined by the current account relation:

$$\frac{s_t f_t F_{t+1}}{p_t} = \frac{s_t F_t}{p_t} + \frac{p_t^H}{p_t} x_t^H - s_t \frac{p_t^F}{p_t} y_t^F - s_t \frac{p_t^e}{p_t} IE_t \tag{31}$$

The net external position F depends on the interest payments and the trade balance: the value of exports x^H minus the imports of final products y^F and energy inputs. Energy acts only as an input in the production process:

$$IE_t = ieA_t K_t^{\alpha} H_t^{1-\alpha} e^{\gamma t} \tag{32}$$

Exports are determined by the price elasticity of foreign demand and by the demand in the rest of the world (ROW):

$$X_t = f(ROW_t, \frac{s_t p_t^F}{p_t^H}) \tag{33}$$

3.5 Market equilibrium

Most of the relations above are derived for an individual household or firm. With the exception of aggregate consumption, presented above, aggregate and individual behavioural equations remain the same, such that the interpretation can switch from the micro to the macro level.

The goods market is in equilibrium if firms' production equals demand by domestic and foreign buyers.

The labour market is in equilibrium if firms' demand for labour equals households' supply. Wages adjust to equilibrate demand and supply. It is assumed that firms use labour inputs of both types of household in fixed proportions, so that the labour supply of the unconstrained households determines the employment outcome and the wage rate following demand shocks. The impact of labour supply elasticity will be discussed later as it will become an important variable in the model outcome.

$$H_t = 1 - L_t \tag{34}$$

So far, the labour market has been considered as a competitive market and this is a very unrealistic hypothesis. In future research it will be replaced with a bargaining.

The demand for money was derived in equation (14). The supply of money follows the following process:

$$M_{t+1} = M_t(1 + gm_t) \tag{35}$$

in which *gm* represents the money supply growth rate. This growth rate can react endogenously to output, inflation or the exchange rate. So different monetary policy reaction functions can be introduced into the model. The effects of monetary policy shocks can be further enhanced by liquidity effects. When such liquidity shocks affect the consumer after his consumption/savings decision is made, they create temporary deviations from the first-order conditions by pushing the nominal interest rate lower.

In the capital market, equilibrium means that government debt is held by domestic investors (assuming that the country is in a positive net foreign asset position) at the market interest rate *R*, and that the net foreign assets are held by investors at the going interest and exchange rates. Both assets are considered to be perfect substitutes, such that the domestic interest rate equals the foreign rate plus expected exchange rate movements (uncovered interest rate parity). The risk premium, present in the first-order conditions, disappears during the linearisation process (certainty equivalence).

3.6 Simulation of public expenditure and tax shocks

The model can be used to simulate the impact of public expenditure cuts and tax increases. As far as possible, the parameters of the model are chosen to reflect the characteristics of the Belgian economy. The structural and technical parameters such as the components of GDP, the wealth composition, the production function, etc. therefore represent a very open economy, with a large public debt and positive net foreign assets vis-à-vis the rest of the world. The calibration of the behavioural parameters, on the other hand, is much more difficult and as the empirical calibration exercise is not yet finished, the obtained impulse responses should not be considered as necessarily representative for the Belgian economy. Therefore, we will confine the analysis to the impact of some of the crucial parameters on the outcome of the simulation results.

Following the discussion in the literature, we examine the impact of the planning horizon and the importance of liquidity constraints, the difference between permanent and transitory shocks, the substitution between private and public consumption, labour supply behaviour, price rigidity and monetary policy reactions. We also briefly present the impact of some other shocks, in particular the impact of a productivity shock as an example of a supply shock and a demand shock (foreign and domestic). These impulse response effects can serve as a benchmark for evaluating the SVAR results in the empirical part of the paper.

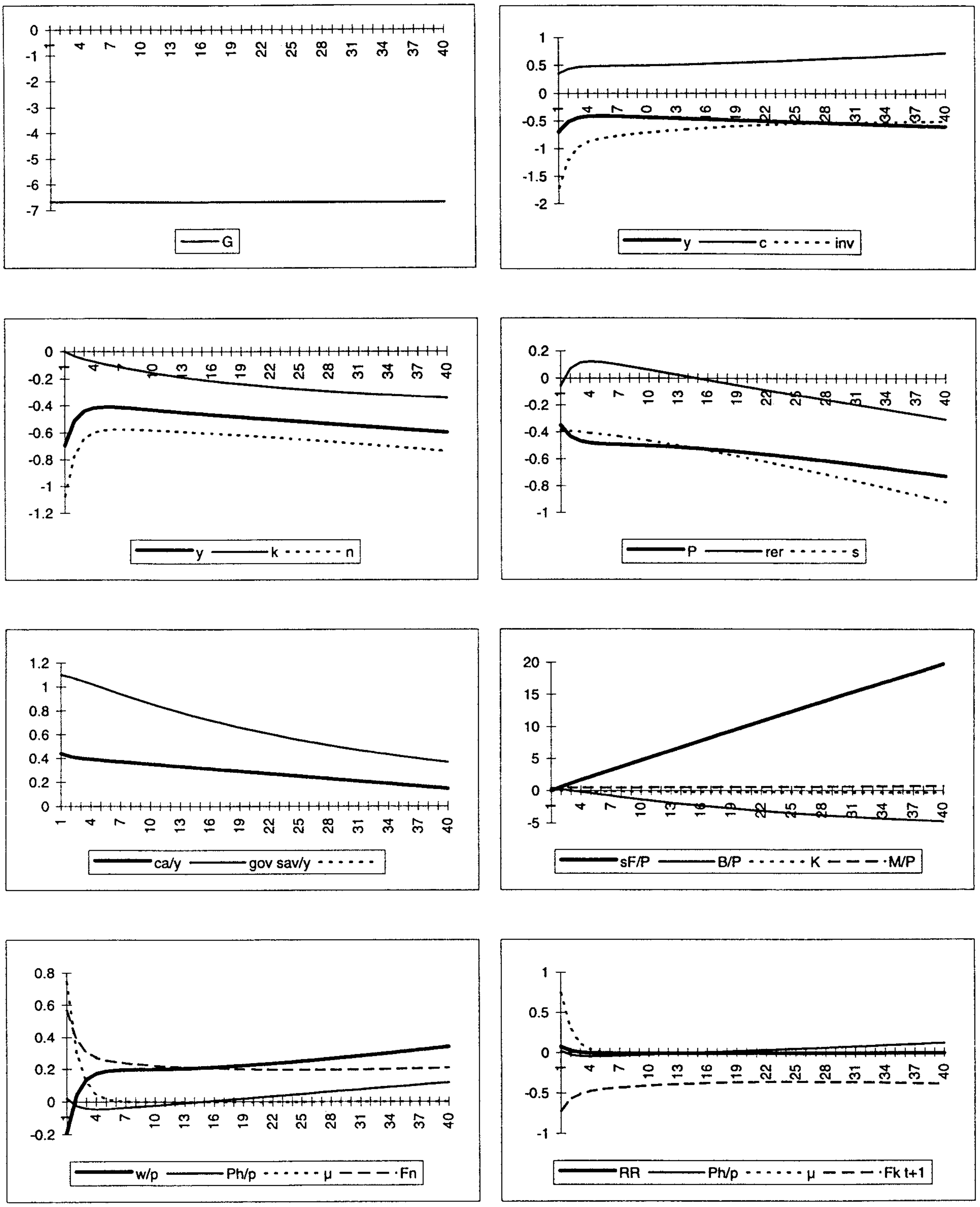
(a) The baseline simulation

We simulate a permanent reduction in government consumption of 1% of GDP and an equivalent increase in lump-sum taxes. These policy actions cause a decline in the public deficit. As public debt decreases below its original level, lump-sum taxes start falling such that the public debt stabilises at a lower level with a multiplier of $1/(g\text{-}RR)$, where g represents the fiscal reaction to the debt level in equation (31) and *RR* is the real interest rate. In the baseline example, this means that the level of public debt is around 5.5% lower in the new steady state. In Figures 2a and 2b, the impulse response outcomes of these simulations are summarised. As the public sector has a large debt service burden, the percentage decrease in public expenditure (6.66%) is higher than the increase in lump-sum taxes (4.5%).

The decrease in public expenditure lowers output (y) by 0.7% and employment (n) by 1.1% during the first quarter. After one year the fall in output is reduced to around 0.4%, but subsequently output converges to a steady-state level of almost 1.5% below the baseline. On the demand side, private consumption reacts positively following the improvement of the private wealth

Figure 2a

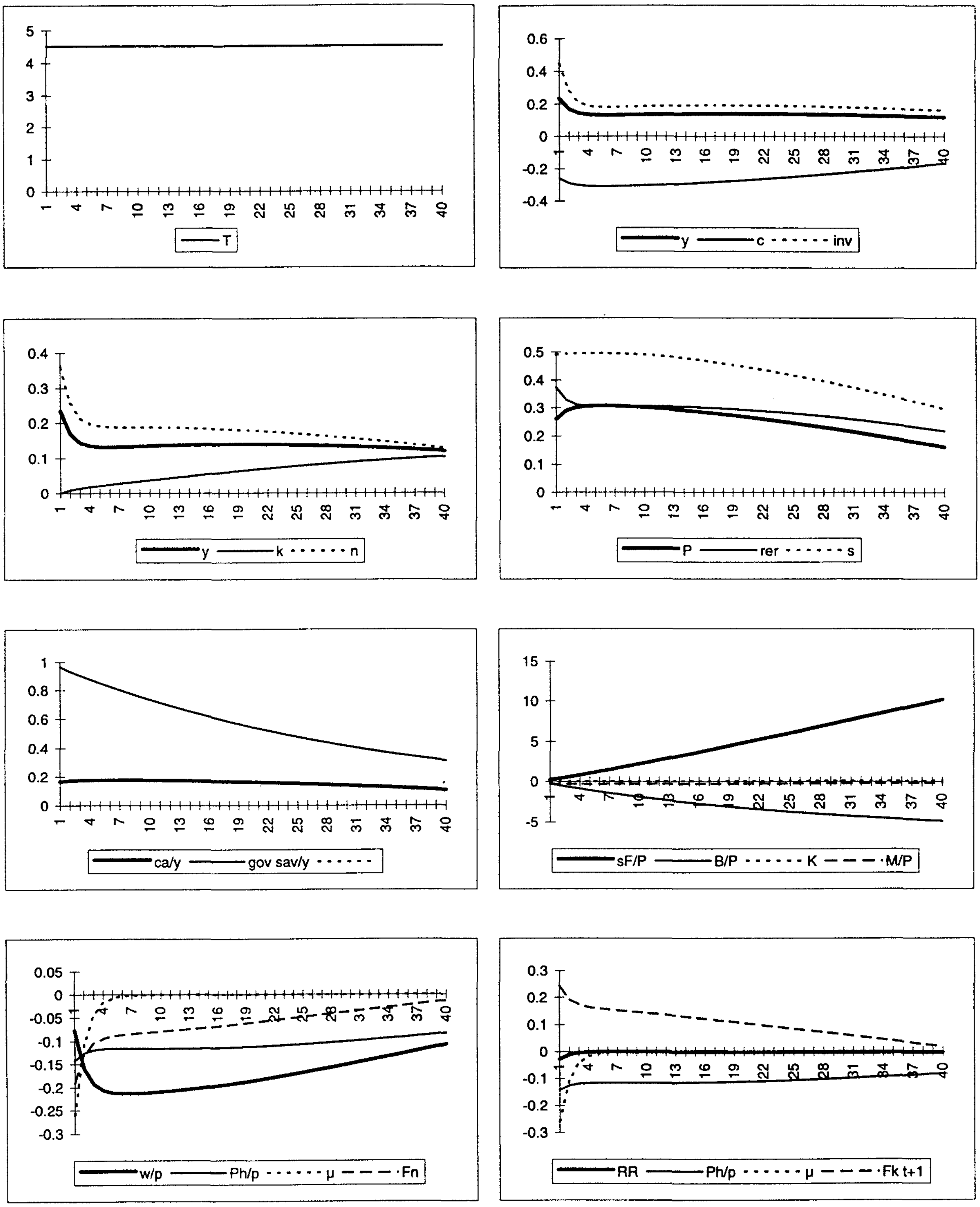
Impulse response of a permanent decrease in government consumption (1% GDP)



Notes: Absolute deviations for the real interest rate *RR*, current account and government savings/GDP ratio; percentage deviations otherwise. An increase in the exchange rate (*s*) represents a depreciation. The real exchange rate (*rer*) is defined as *spf/ph*.

Figure 2b

Impulse response of a permanent decrease in the exogenous taxes (1% GDP)



Notes: Absolute deviations for the real interest rate *RR*, current account and government savings/GDP ratio; percentage deviations otherwise. An increase in the exchange rate (*s*) represents a depreciation. The real exchange rate (*rer*) is defined as *spf/ph*.

position. As consumers have a finite horizon (50 quarters or 12.5 years in the baseline scenario) and taxes decrease gradually over time in line with the debt reduction, private wealth and consumption show a further increase over time as future generations profit more from the consolidation process. Investments (*inv*) decline strongly in the short run and further depress aggregate demand. Increasing consumption goes hand in hand with an increasing demand for leisure or a declining labour supply. Together with the lower capital accumulation (k), this explains the gradual decline of output in the long run.

In the baseline simulation, nominal money supply is kept constant. As money demand is specified in terms of consumption, the aggregate demand price deflator (p) has to decline in order to equalise money demand and supply. The nominal interest rate does not change as the higher real rate is compensated for by lower expected inflation. With a constant nominal interest rate, the exchange rate (s) has to jump directly towards its new equilibrium path. During the first few years, the nominal appreciation does not offset the decline in the price level so that the real exchange rate (*rer*) depreciates. The subsequent weak real exchange rate appreciation is in line with the lower real interest rate.

The real depreciation during the first few years helps to offset the negative domestic demand shock. It is, however, limited by the rigidity of prices and the corresponding downward reaction of aggregate supply. The current account (*ca*) strongly improves through the decrease in imports and the improvement in competitiveness. As the net foreign asset position (sF/p) improves over time, the exchange rate starts to appreciate. In the new steady state, equilibrium in the current account requires a real appreciation and a trade deficit to compensate for higher interest income from abroad. This current account and real exchange rate behaviour is in line with the growing divergence between private consumption and output.

The composition of financial wealth shifts away from public debt and equity towards foreign assets. As the decrease in public debt finds its counterpart in the lower present value of future taxes, and therefore higher human wealth, the accumulation of foreign assets reflects, to a large extent, a net increase in private wealth. This evolution is in line with the consumption and labour supply behaviour of the households.

The impulse response effects on the supply variables are important in explaining the short-run output costs. Marginal costs are a function of labour productivity (Fn), the real wage (w/p) and the relative output price (ph/p). At a lower output level, marginal costs decrease as labour productivity increases and real wages in terms of output prices decline. Prices follow the lower production cost only gradually as firms temporarily increase their mark-ups. Higher mark-ups shift the demand for labour and the supply by firms downwards. Higher mark-ups and lower marginal productivity of capital also explain the strong decline in Tobin's Q and investment. Together, these negative supply reactions of firms explain the high short-run output costs of a demand shock in this new-Keynesian framework.

A fiscal consolidation through a lump-sum tax increase or a decrease in transfers has a very different impact on the economy (Figure 2b). As consumers have a finite horizon, the burden of public debt shifts from the future to the present generation, and private consumption will decrease. But aggregate demand increases as investment and, especially, net exports rise strongly. So there is a shift in the use of resources away from consumption towards capital accumulation, both domestically and abroad in the form of net foreign assets.

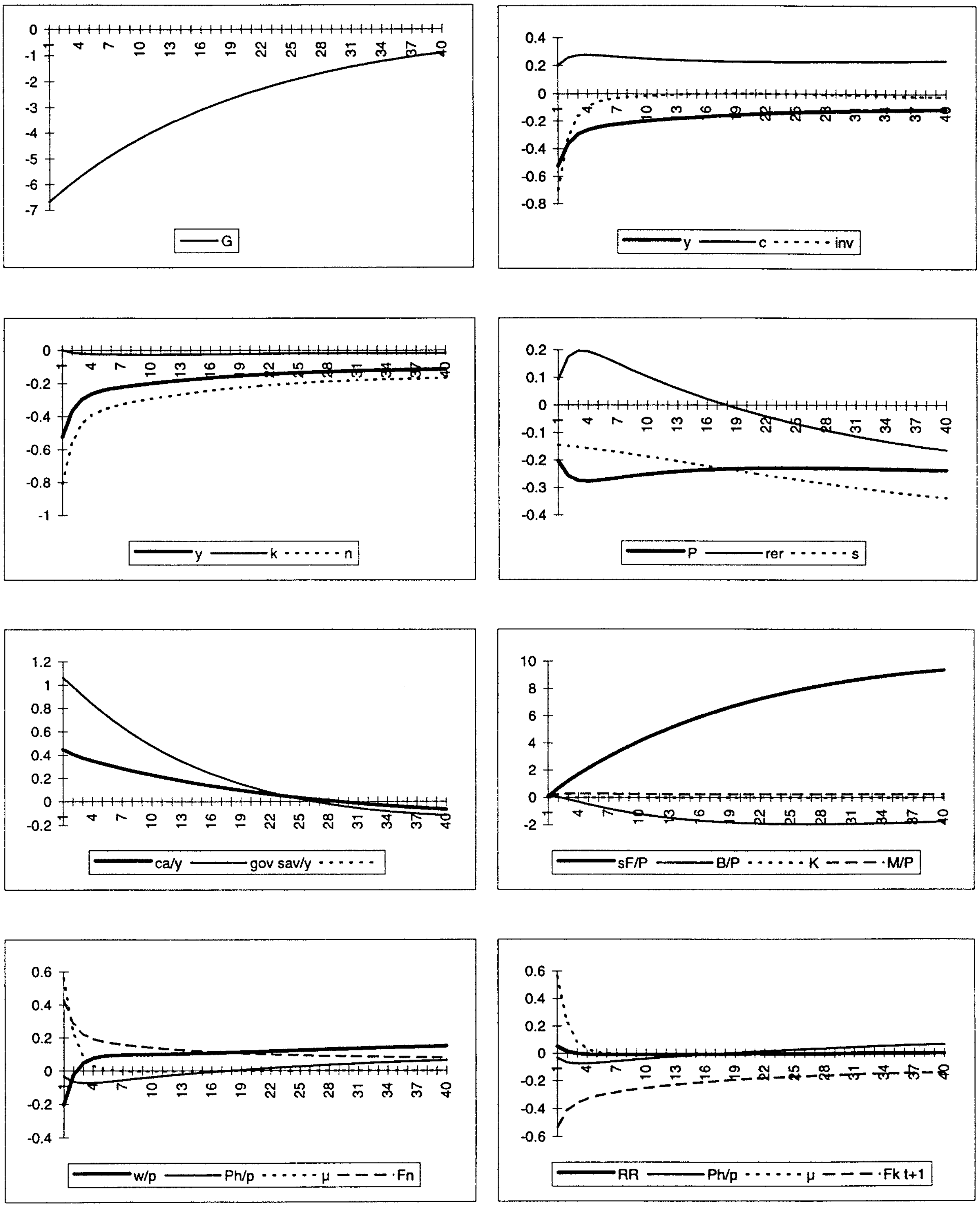
The positive net export evolution is possible because there is a strong real depreciation of the exchange rate, which compensates for the higher demand for imports following the increase in aggregate demand. Over time the exchange rate will appreciate again so that in the new steady state net exports will turn negative and compensate for the higher capital income.

The increase in total demand is accompanied by an increase in prices but a decrease in the mark-up. Temporarily lower mark-ups push labour demand, investment and production up further. This extra supply of domestic goods, in the short run, enhances the downward pressure on the price of

Figure 3

Impulse response of a transitory decrease in government consumption (1% GDP)

Persistence $\rho = 0.95$



Notes: Absolute deviations for the real interest rate *RR*, current account and government savings/GDP ratio; percentage deviations otherwise. An increase in the exchange rate (*s*) represents a depreciation. The real exchange rate (*rer*) is defined as *spf/ph*.

the domestic good in terms of the foreign good. Over a longer period, there is a decrease in the marginal productivity of labour and a decline in the relative output price of the home good, but these negative influences on the demand for labour are compensated for by a lower real wage.

(b) Transitory versus permanent shocks

In the literature, some authors emphasise that permanent public spending cuts should be less costly than transitory measures. Private wealth and expectations should react more positively to permanent measures. We therefore compare the impact of a transitory spending cut with the results of a permanent spending cut as described in the baseline projection (Figure 3).

A transitory spending cut has a smaller wealth effect, so consumption increases less than it would in response to a permanent measure. This reflects the basic argument behind the original expansionary fiscal contraction. But of course our model does not contain the non-linear effects that can further strengthen the normal wealth effect via a shift in expectations on future fiscal policy.

However, in our model this lower increase in private consumption does not translate into a higher short-run output cost for a strict fiscal policy. On the contrary, the short-run output costs of a transitory shock are lower because investment declines less and because there is a stronger real depreciation and, therefore, a better performance by exports.

(c) The impact of the planning horizon and liquidity constraints

Figure 4 shows the result for a permanent spending cut with a lower expected life for households: five years instead of 12.5 in the baseline simulation. The length of the horizon is an important determinant for the strength of the wealth effect. By increasing the probability of death, the discount rate of households for future income and taxes increases. The decrease in future taxes, following the fiscal consolidation, therefore receives a much lower weighting in the calculations of the households. Wealth increases less, and private consumption will also increase less during the first period of the simulation. As taxes are effectively lowered later in the process, consumption tends towards the same level as in the baseline simulation.

The smaller impact on consumption is again compensated for by more investment and, in particular, by a stronger real depreciation and, therefore, higher exports. Unemployment and output are also higher because of the improvement in the supply conditions: real wages decline relative to the baseline simulation with a similar productivity. This result illustrates that for a small open economy the impact of the fiscal programme on competitiveness and foreign demand can be more important in determining the output costs than the impact on domestic demand.

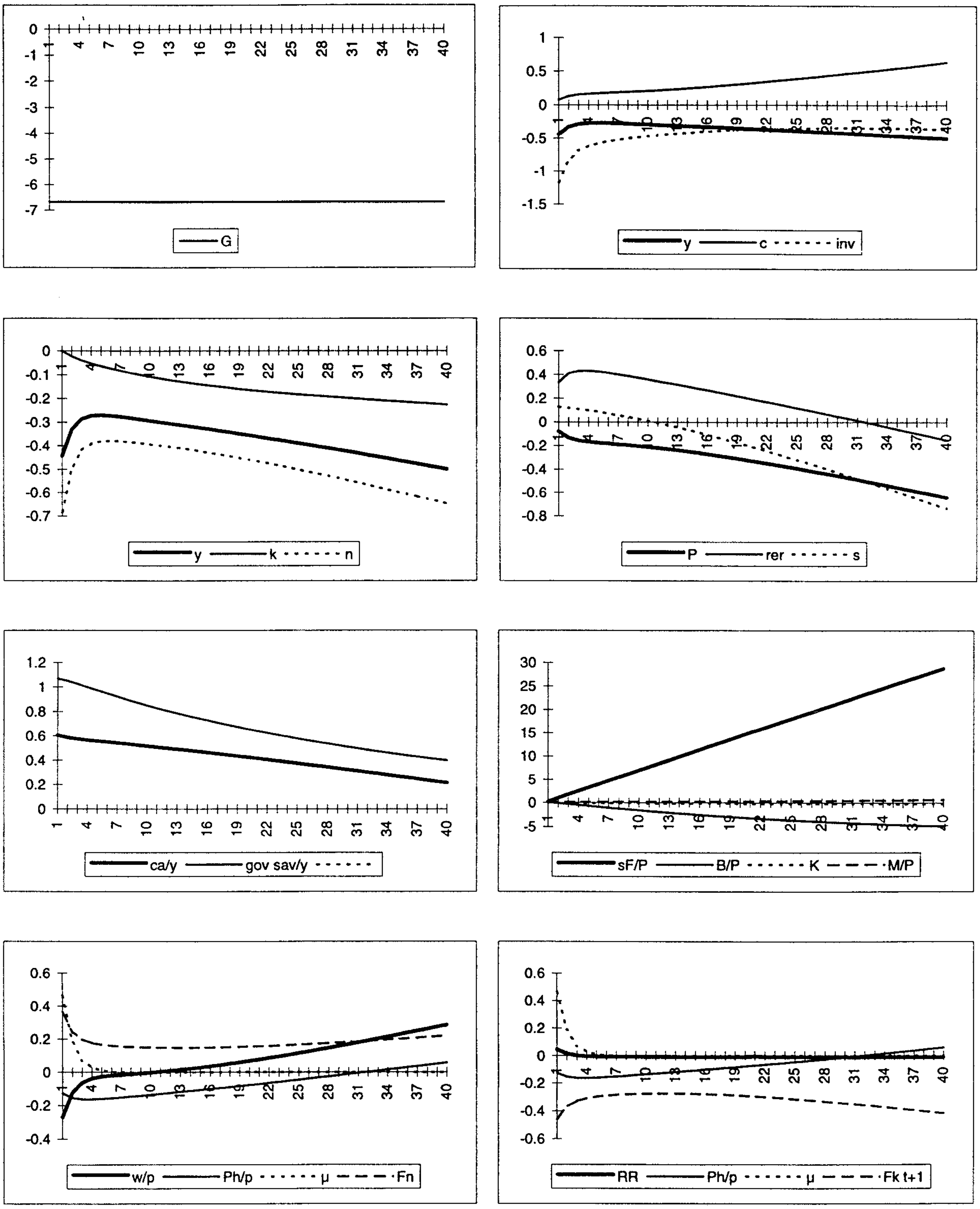
The relevance of this simulation is probably also important as it can also capture the case of higher uncertainty about future fiscal policy. If the fiscal consolidation process is not considered credible by households, they will also discount future tax promises at a higher discount rate. The lower domestic demand that results in this case could be offset by a stronger exchange rate depreciation and lower real wages, which stimulate foreign demand.

The impact of liquidity constraints on a certain proportion of households has a more complex effect. Lower output and labour income decrease the income and consumption of the constrained households. Consumption of the unconstrained households will, however, go up, but this result depends on the specification of the production function and the corresponding income distribution. The output cost increases in this scenario as the labour supply is lower and real wages higher, but these effects are rather small compared with the influence of the horizon length.

If the proportion of liquidity-constrained households is further increased, the dynamic properties of the model change and the solution path is no longer uniquely determined. Sunspots and self-fulfilling expectations allow a large diversity of outcomes in this case, so that a general conclusion can no longer be drawn.

Figure 4

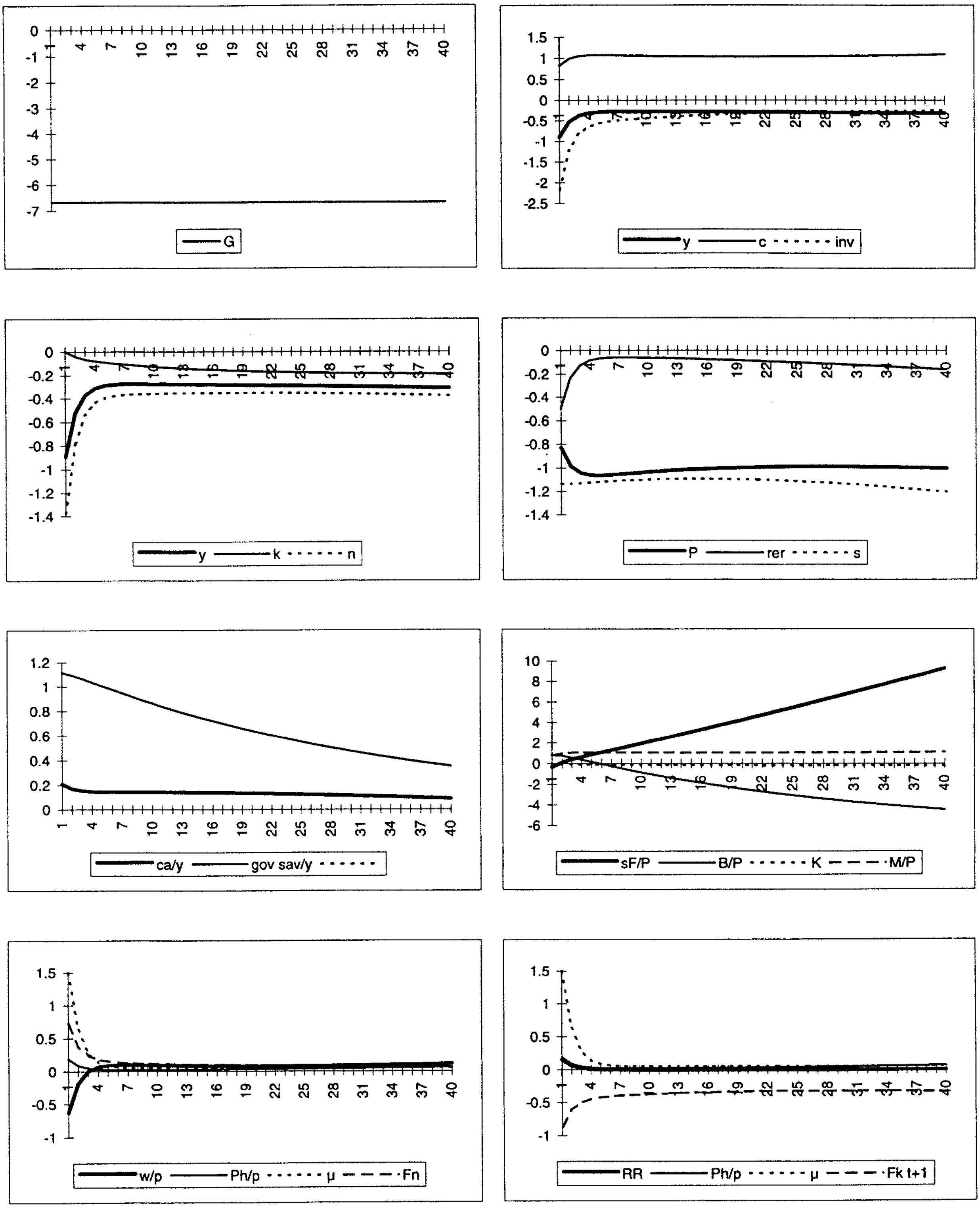
Impulse response of a transitory decrease in government consumption (1% GDP): scenario with a shorter expected life



Notes: Absolute deviations for the real interest rate *RR*, current account and government savings/GDP ratio; percentage deviations otherwise. An increase in the exchange rate (*s*) represents a depreciation. The real exchange rate (*rer*) is defined as *spf/ph*.

Figure 5

Impulse response of a permanent decrease in government consumption (1% GDP): scenario with substitution between private and public consumption (θ = 0.5)



Notes: Absolute deviations for the real interest rate *RR*, current account and government savings/GDP ratio; percentage deviations otherwise. An increase in the exchange rate (*s*) represents a depreciation. The real exchange rate (*rer*) is defined as *spf/ph*.

(d) Substitution between private and public consumption

As an alternative scenario, we consider the case where public consumption is a substitute for private consumption (Figure 5). This affects the model via two channels. Public consumption is added to the utility function together with private consumption (with a positive coefficient in the case of substitution) and the present value of future public consumption (with the same coefficient) is added to the disposable wealth of the households.

If the government reduces the supply of public goods, this will increase the private consumption demand in the case of substitution. But a decline in the supply of public goods also directly decreases the wealth constraint: the gain from lower future taxes is offset by the fall in the present value of future public goods. In the extreme case, where public goods are a perfect substitute for private consumption, there would be no impact on the model. With less-than-perfect substitution, private consumption will increase more than in the baseline simulation, and the output costs will be lower.

(e) Labour supply behaviour

With a utility function that is linear in leisure, aggregate labour supply becomes infinitely elastic (Hansen (1985)). In this scenario both private consumption and output turn out to be lower (Figure 6). Firms can adjust supply and employment more easily following the decline in aggregate demand. As there is no downward pressure on real wages from the decline in employment in this scenario, real wages increase more in line with the higher marginal productivity of labour. Both factors limit the real depreciation of the currency, so that exports further depress aggregate demand. The current account remains positive as import demand also decreases.

This result illustrates the importance of the supply-side reaction in determining the outcome of a fiscal adjustment programme, especially for an open economy. In our simple model of the labour market, this effect depends only on the elasticity of the supply of workers. In reality, this effect will be much more complex. A fiscal adjustment programme, by cutting public employment, reducing public sector wages and lowering social security payments, is likely to lower the reservation wage of workers. This effect reduces the bargaining power of labour unions in the wage negotiations. In such a context, a strict fiscal adjustment programme is likely to also create a positive supply shock. By lowering real wage costs and increasing profitability, firms will be stimulated to decrease prices and increase production, in particular by boosting exports. By decreasing the labour supply elasticity in our model, we can generate a similar effect. But further development of the labour market block is necessary to obtain a realistic representation of the supply-side behaviour and to analyse the impact of these complications on the rest of the model.

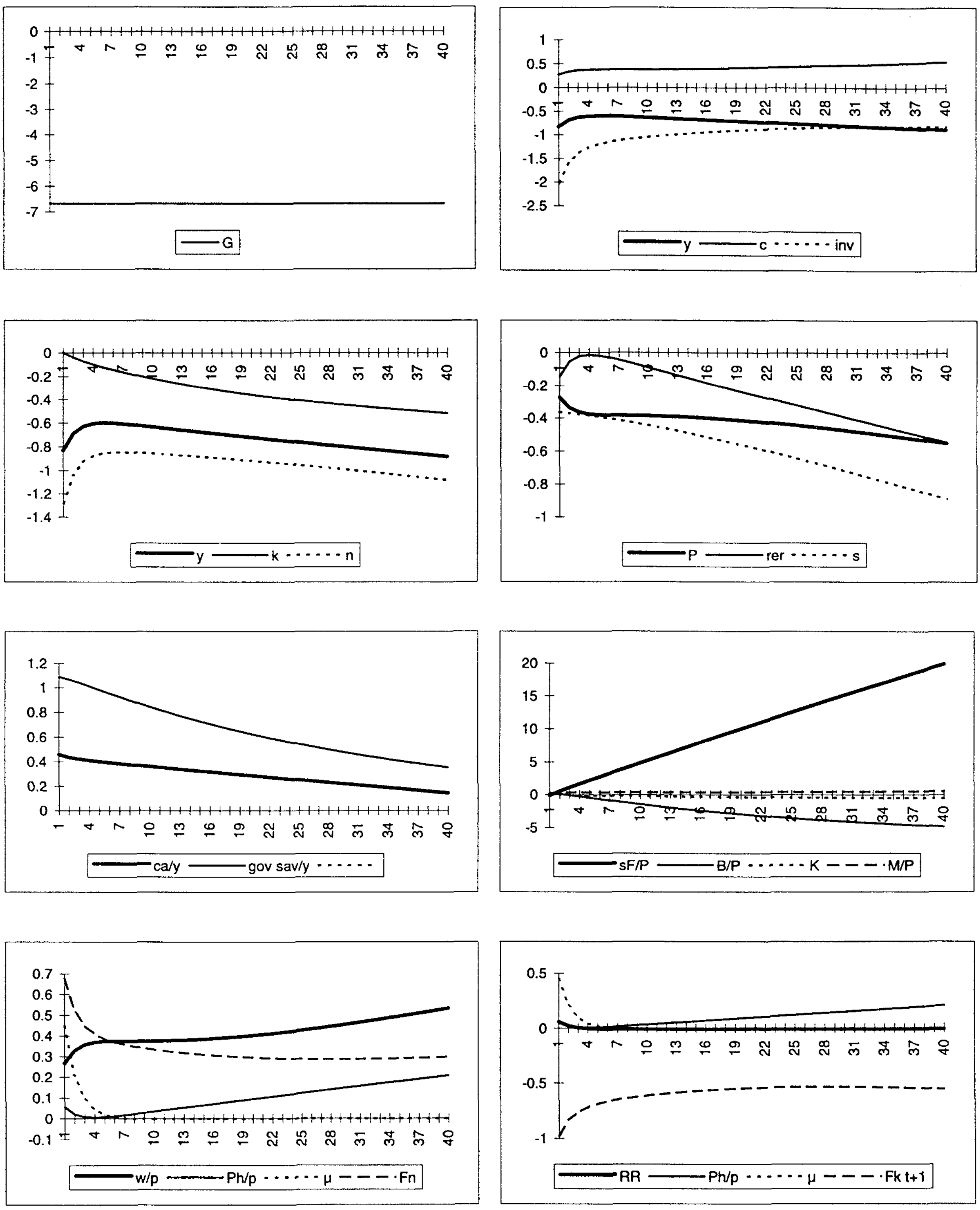
(f) Degree of price rigidity

The degree of price rigidity is the crucial variable in the determination of the short-run output costs of a negative demand shock. In Figure 7, we present the simulation results for the model where the adjustment speed of prices is decreased from 0.4 in the baseline to 0.1.

In this scenario, a negative demand shock results in higher and more persistent mark-ups. As firms pursue higher profit margins they limit employment, investment and output. Since prices do not follow the declining marginal costs, the real exchange rate will appreciate as the relative price decline of the domestic good is smaller than the nominal appreciation. Despite the fall in real wages, the competitive position worsens because firms are unwilling to pass on lower costs to output prices. So exports will also decline, further aggravating the decrease in aggregate demand. Once the price adjustment process is finished, the results converge to the same long-run effect as in the baseline simulation.

Figure 6

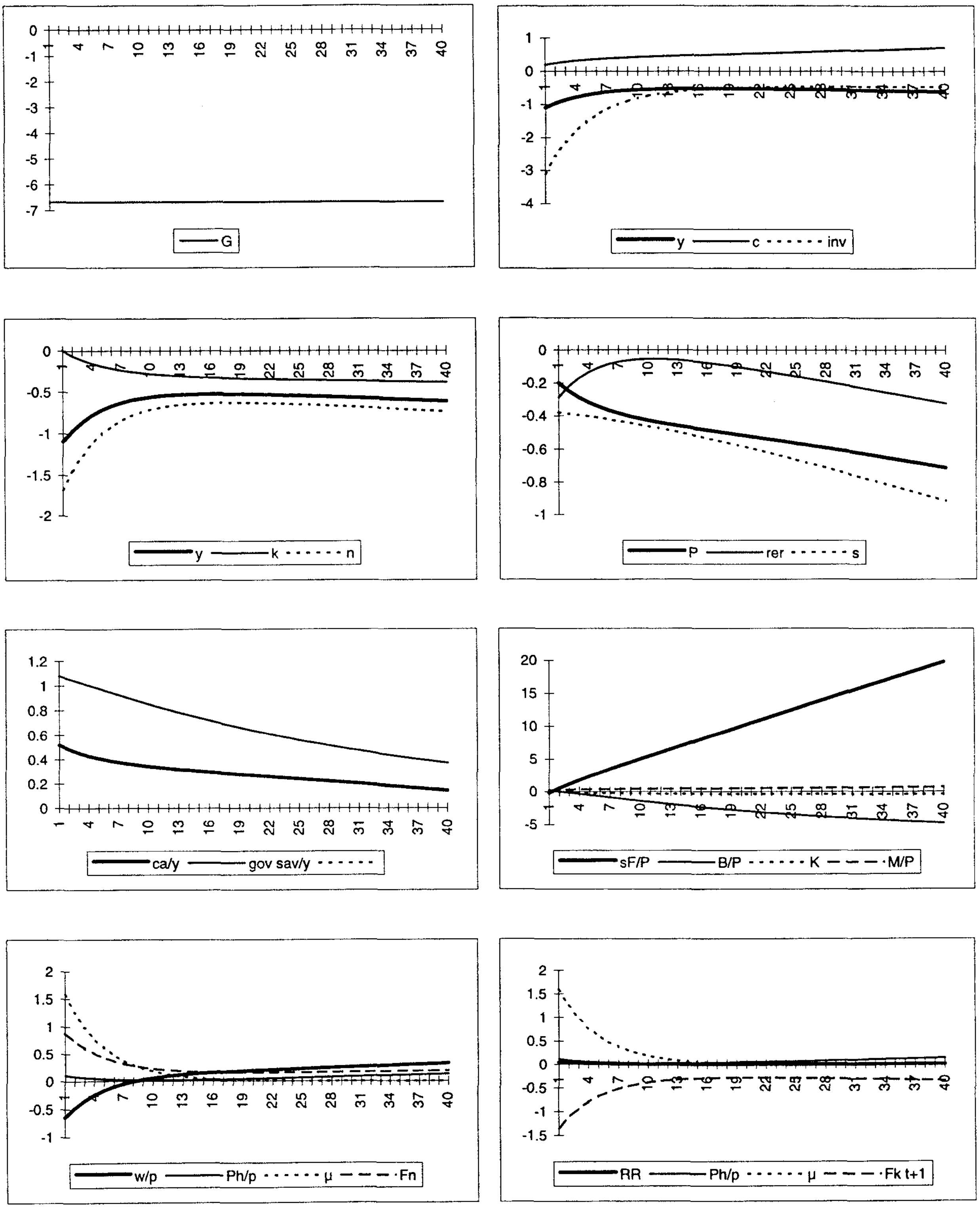
Impulse response of a permanent decrease in government consumption (1% GDP): scenario with infinite elastic labour supply



Notes: Absolute deviations for the real interest rate *RR*, current account and government savings/GDP ratio; percentage deviations otherwise. An increase in the exchange rate (*s*) represents a depreciation. The real exchange rate (*rer*) is defined as *spf/ph*.

Figure 7

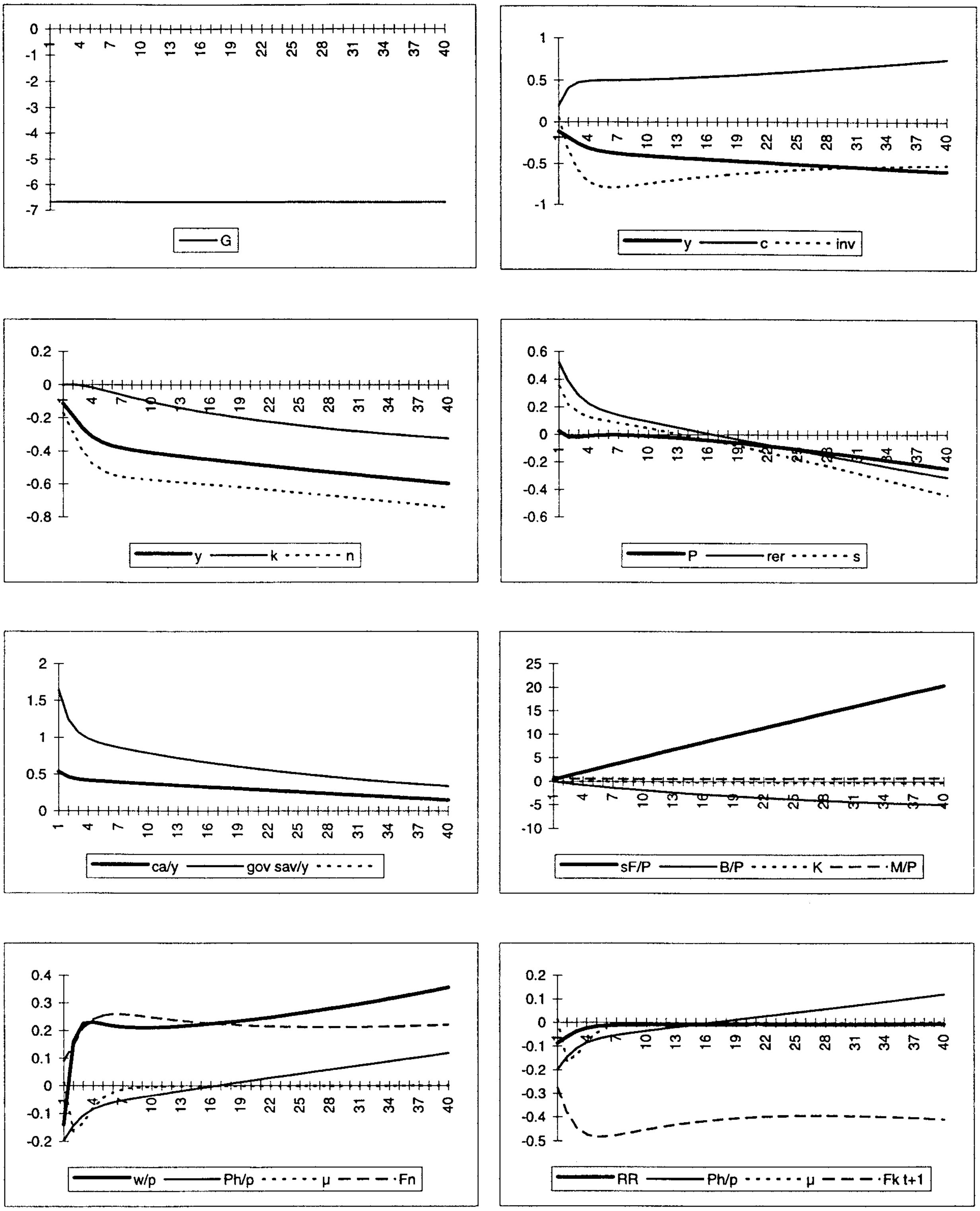
Impulse response of a permanent decrease in government consumption (1% GDP): scenario with higher price rigidity (ζ = 0.9)



Notes: Absolute deviations for the real interest rate *RR*, current account and government savings/GDP ratio; percentage deviations otherwise. An increase in the exchange rate (*s*) represents a depreciation. The real exchange rate (*rer*) is defined as *spf*/*ph*.

Figure 8

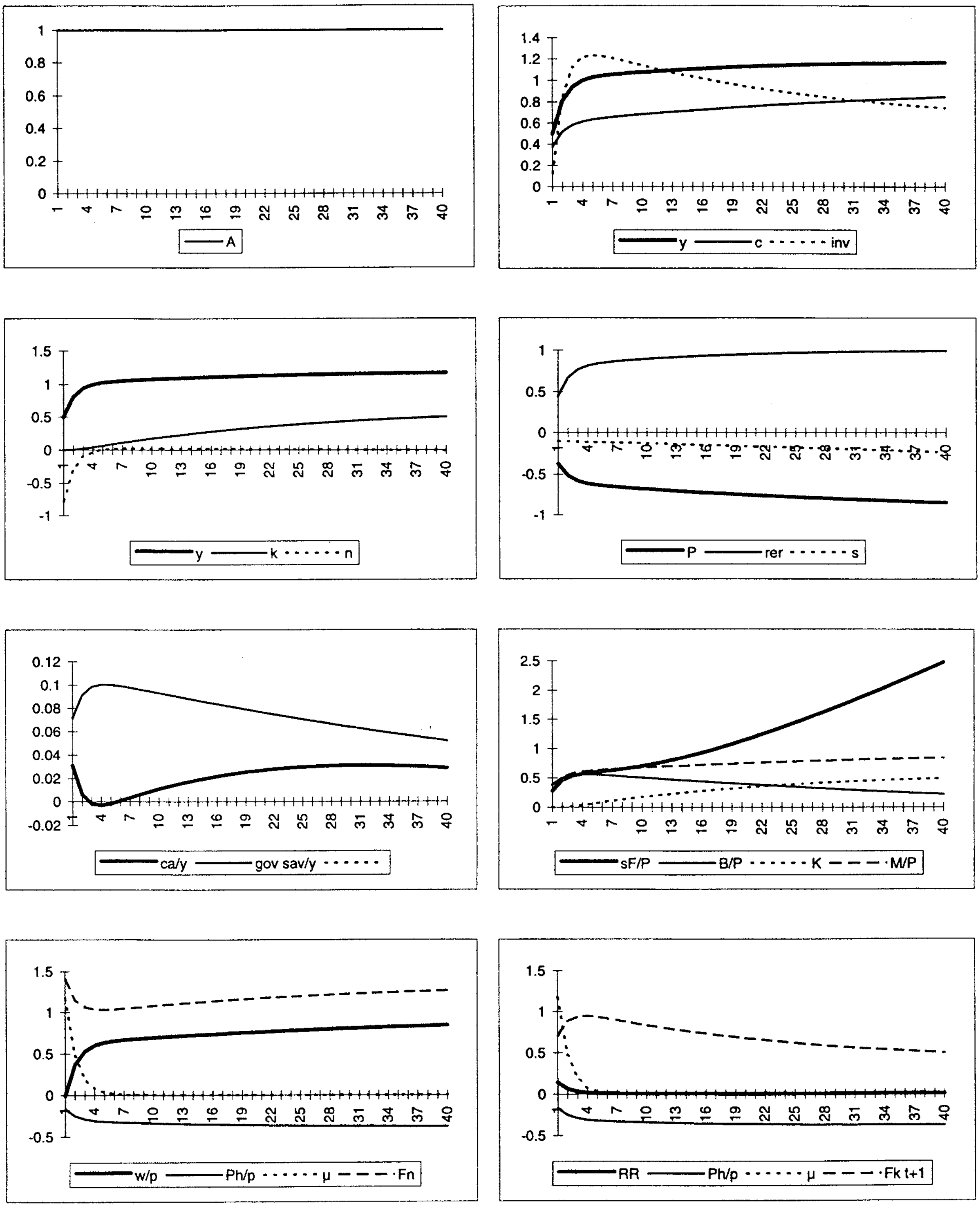
Impulse response of a permanent decrease in government consumption (1% GDP): expansive monetary policy to stabilise the price level



Notes: Absolute deviations for the real interest rate *RR*, current account and government savings/GDP ratio; percentage deviations otherwise. An increase in the exchange rate (*s*) represents a depreciation. The real exchange rate (*rer*) is defined as *spf/ph*.

Figure 9

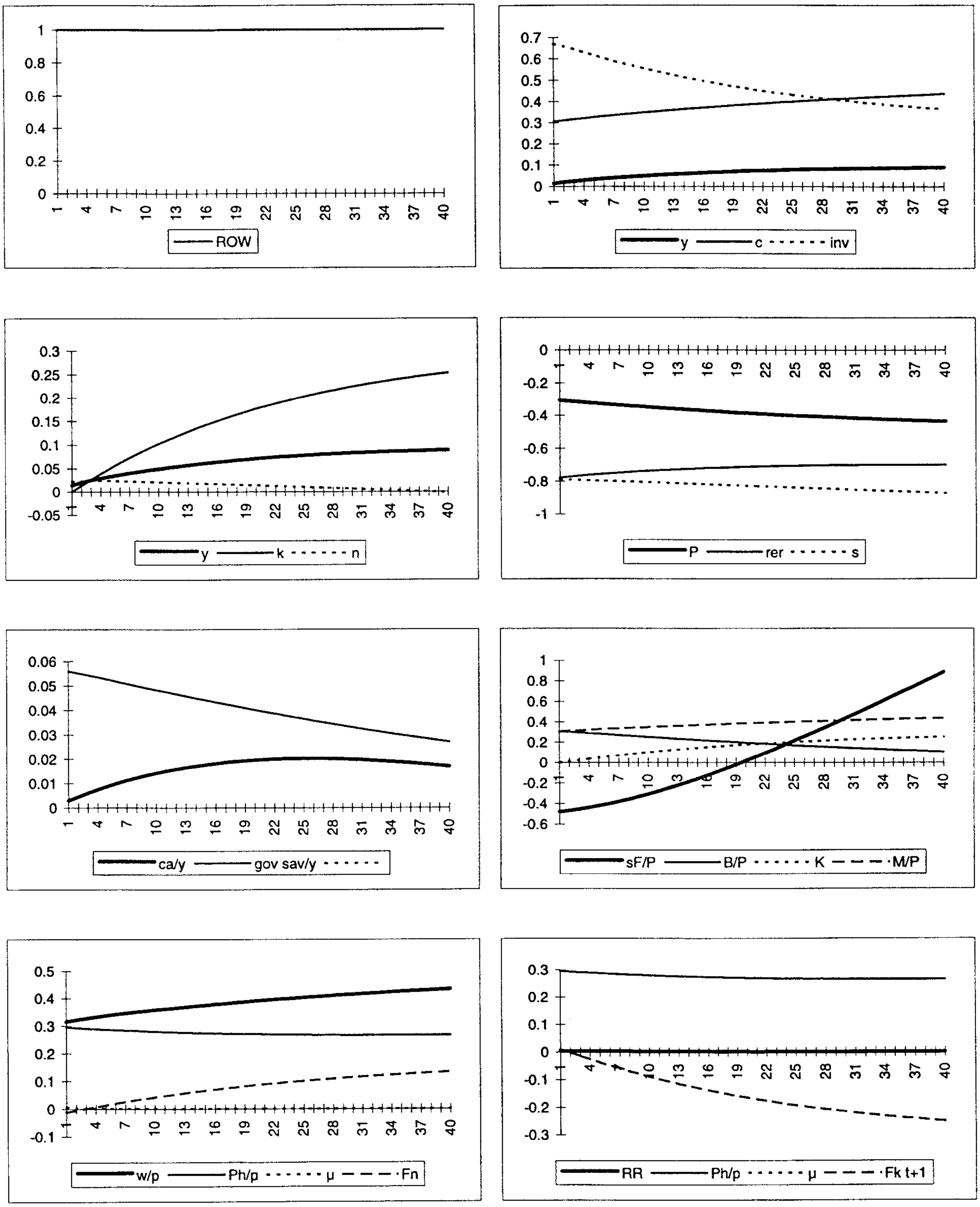
Impulse response of a permanent productivity shock



Notes: Absolute deviations for the real interest rate *RR*, current account and government savings/GDP ratio; percentage deviations otherwise. An increase in the exchange rate (*s*) represents a depreciation. The real exchange rate (*rer*) is defined as *spf/ph*.

Figure 10

Impulse response of a permanent increase in foreign demand



Notes: Absolute deviations for the real interest rate *RR*, current account and government savings/GDP ratio; percentage deviations otherwise. An increase in the exchange rate (*s*) represents a depreciation. The real exchange rate (*rer*) is defined as *spf/ph*.

(g) *Monetary policy behaviour*

In the baseline simulation, a cut in public expenditure produces a decline in the price level and an appreciation of the nominal exchange rate. This result is obtained with a fixed nominal money supply rule. Under this behaviour, the real interest rate increases during the price adjustment process, which shows the rather restrictive character of the policy.

If monetary policy targeted a fixed nominal exchange rate or inflation, it would react more expansively. A combination of a restrictive fiscal policy and an expansionary monetary policy lowers significantly the short-run real output costs of a fiscal consolidation.

In Figure 8, the results of a spending cut are shown for the hypothesis that monetary policy tries to stabilise the price level, and will therefore react with a loose stance. A policy aimed at stabilising the nominal exchange rate should be somewhat less expansive, but the result goes in the same direction. The stabilisation of the price level is obtained through a depreciation of the nominal exchange rate, so that lower domestic output prices are offset by higher import prices.

Such a monetary policy reaction is able to neutralise a large portion of the short-run output costs. The real depreciation stimulates exports and the decrease in interest rates supports investment. As output prices or marginal costs are more stable, the mark-up also remains stable (or declines temporarily), so that one can also prevent negative supply reactions of firms. But these expansionary effects of monetary policy are short-lived, and after one or two years the economy is back on the same dynamic path.

Finally, in Figures 9 and 10 the impulse response to a productivity shock, as an example of a favourable supply shock, and an increase in demand in the rest of the world for the domestic good, as an example of a positive demand shock, are presented. These outcomes can serve as a benchmark for evaluating the effects of the SVAR estimation results in the next section.

4. Structural VAR estimation results

The SVAR approach is an appropriate technique for estimating the impulse response to public spending shocks. It allows us to identify the specific impact of government expenditure on economic growth, after correcting for other macroeconomic shocks. This separation of different shocks is important as many studies on the effects of consolidation report the problem of a number of different shocks occurring simultaneously. It is therefore crucial to isolate the specific role of public spending measures in the observed growth process.

In our exercise, we use a very simple model with three variables: GDP growth, inflation and growth in government expenditure. The structural identification restrictions determine three types of shock: supply shocks, demand shocks and public spending shocks. Small open economies depend heavily on developments elsewhere in the world; we therefore include foreign growth, inflation and short-term interest rates as exogenous variables in the model. The three structural shocks explain the remaining fluctuations of domestic origin. Of course, the reduction of the observed fluctuations to three shocks and three foreign variables is a strong simplification of reality, and the results of the exercise should therefore be analysed critically and considered as rough indications.

The model is estimated for three European countries: Belgium, Ireland and Denmark, using annual data over the period 1964-96.All data are taken from the EEC Annual Macro Economic Data Base. Real government expenditure is represented by the (national accounts) series for public consumption in constant prices. Inflation is measured by the log change in the GDP price deflator and growth by the change in the log of GDP at constant market prices. German GDP growth, inflation and three-month interest rates are used as exogenous variables.

Following the discussion of the theoretical model in the previous section (but which is also generally accepted in the SVAR literature), demand shocks, including monetary policy shocks, are assumed to have no important effects on output in the long run. The long-run equilibrium output depends on foreign shocks, supply shocks and, as this is the central topic of the paper, possibly also on public spending shocks. As public spending directly affects the budget constraints of households, it is more likely to have long-run effects on output than demand shocks stemming from other sources. This theoretical restriction on the long-run outcome of demand shocks is combined with specific restrictions to distinguish public spending shocks from the other types of disturbance. Two variants were used. In one version, supply and demand shocks do not affect public spending during the period in which the shock occurs, that is, all "innovations" in public spending are considered as public spending shocks. In the second variant, we use long-run restrictions and assume that government spending in the long run depends only on policy decisions and is independent of supply or demand shocks. This second version is used to check the sensitivity of the results to the specific form of the restrictions used. But it also changes somewhat the interpretation of the results as it only looks for permanent changes in public spending. Transitory shocks are excluded and this may possibly change the impact on output and inflation.

The same exercise was performed using real exchange rate changes vis-à-vis the Deutsche mark (defined with the GDP deflator) instead of inflation. By comparing the results for inflation and the real exchange rate, one should get an indication of how the nominal exchange rate and monetary policy react to public spending shocks. If the real exchange rate moves by much less than inflation, this means that the nominal exchange rate movements were compensating for the relative price developments. So if spending cuts put downward pressure on the domestic price level and inflation, the real exchange rate should depreciate, unless the nominal exchange rate was appreciating strongly. Such an appreciation is most likely if monetary policy is independent from fiscal shocks, so that interest rates do not decline strongly following the fiscal adjustment and the fall in inflation. This monetary policy reaction should increase the real impact of spending shocks, and the negative pressure on prices will also be reinforced by the nominal appreciation. On the other hand, if monetary policy is rather expansive following a restrictive spending policy, the real exchange rate will depreciate more than the price level and fiscal policy will have smaller short-run multiplier effects on output, with less reaction in prices.

To ensure that the economic interpretations of the shocks make sense, the historical series of the three shocks are used as explanatory variables in simple autoregressive equations for a set of macroeconomic variables related to supply, demand and public finance conditions. These regressions should indicate whether the structural error series are indeed correlated with innovations in the macroeconomic variables they are supposed to summarise (Table 1).

In all three countries, the supply shock is significantly positively correlated with the innovations in total factor productivity and labour productivity. In Belgium, the supply shock is also significantly negatively correlated with the change in the tax burden as measured by income taxes and social security contributions as a proportion of total compensation of employees. In Denmark, a negative correlation with the trade balance is found, probably explained by a strong import content of investment.

The results for the demand shock are more diverse: in Belgium, it is significantly positively correlated with final demand (total and national), but also with the public deficit. This last effect can be interpreted as the result of stronger domestic demand, nominal growth and income. In Ireland, the demand shocks are correlated negatively with innovations in taxes, while for Denmark no significant relations are found. Government spending shocks are negatively correlated with public deficits in Ireland, but not in Denmark, where a strong correlation is found with taxes. In Denmark, government consumption is also positively correlated with productivity and final demand, but negatively with the trade balance.

So while the results for the supply and public spending shocks are acceptable, the demand shocks are less easy to identify. As demand shocks represent a diversity of disturbances that affect the economy in the short run, it is probably acceptable that a stable relation is not shown with any of the individual variables tested over the whole period.

Table 1

Marginal significance of the shock variables in autoregressive equations for a set of macroeconomic variables

	Belgium						**Ireland**						**Denmark**					
	Supply shock		**Demand shock**		**Public spending shock**		**Supply shock**		**Demand shock**		**Public spending shock**		**Supply shock**		**Demand shock**		**Public spending shock**	
GDP (*cte* prices)	0.78	[2.36]	0.82	[2.52]	-0.14	[-0.40]	1.89	[7.53]	0.31	[0.72]	0.96	[2.43]	1.09	[3.16]	0.14	[0.34]	0.79	[2.10]
GDP deflator	-0.50	[-1.92]	0.75	[3.19]	-0.13	[-0.47]	-1.20	[-1.92]	2.15	[4.03]	-0.31	[-0.46]	-0.40	[-1.67]	0.81	[4.04]	-0.19	[-0.74]
Real exchange rate	0.39	[0.66]	-0.33	[-0.53]	-0.10	[-0.16]	2.09	[2.21]	-2.09	[-2.24]	-0.41	[-42.00]	0.15	[0.27]	-1.10	[-2.26]	0.67	[1.33]
Public expenditures (*cte* prices)	0.00	[0.00]	0.00	[0.00]	1.16	[6.97]	0.00	[0.00]	0.00	[0.00]	2.06	[5.54]	0.00	[0.00]	0.00	[0.00]	1.33	[8.94]
Total factor productivity	0.65	[2.42]	0.53	[1.89]	-0.12	[-0.42]	1.41	[5.55]	0.42	[1.17]	0.59	[1.71]	0.79	[3.06]	-0.11	[-0.36]	0.58	[2.10]
Real compensation per person	0.95	[3.22]	-0.14	[-0.40]	-0.04	[-0.10]	0.68	[1.67]	-0.59	[-1.33]	0.50	[1.26]	0.11	[0.33]	-0.19	[-0.64]	0.45	[1.36]
Labour productivity	0.67	[2.55]	0.36	[1.27]	-0.16	[-0.54]	1.08	[4.06]	0.56	[1.74]	0.58	[1.82]	0.64	[2.36]	-0.07	[-0.24]	0.59	[2.15]
Tax burden on labour income	-14.00	[-3.40]	0.01	[0.02]	-0.02	[-0.36]	1.55	[0.04]	-42.90	[-1.38]	-0.33	[-1.00]	0.55	[1.03]	-0.22	[0.38]	1.09	[2.36]
National final demand (*cte* prices)	0.41	[1.28]	0.85	[2.89]	-0.05	[-0.14]	0.96	[1.71]	0.22	[0.38]	0.94	[1.71]	1.81	[3.63]	0.50	[0.84]	1.44	[2.67]
Exports (*cte* prices)	0.93	[1.14]	1.09	[1.32]	0.28	[0.33]	1.34	[1.87]	-0.78	[-1.01]	0.65	[0.78]	-0.54	[-1.09]	-0.25	[-0.49]	0.13	[0.25]
Tax receipts	-0.70	[-0.86]	0.50	[0.61]	1.39	[1.60]	-1.22	[-1.12]	-1.27	[-1.91]	-0.69	[-0.90]	1.03	[1.22]	-0.42	[-0.46]	1.78	[2.94]
Total receipts	-0.59	[-1.49]	0.42	[1.04]	0.21	[0.52]	-0.58	[-0.97]	-0.79	[-1.29]	-0.69	[-0.61]	0.55	[0.56]	-0.66	[-0.66]	2.11	[2.49]
Public transfers	-0.75	[-1.08]	-0.84	[-1.41]	0.39	[0.55]	0.05	[0.04]	0.08	[0.96]	1.81	[1.75]	-1.44	[-2.24]	0.44	[0.61]	0.00	[0.00]
Public deficit	0.12	[0.49]	0.58	[2.75]	-0.21	[-0.86]	-0.35	[-1.07]	-0.39	[-1.34]	-0.72	[-2.65]	0.64	[2.04]	-0.17	[-0.50]	0.21	[0.65]
Export/import ratio (*cte* prices)	0.56	[1.73]	-0.25	[-0.68]	0.09	[0.26]	0.34	[0.19]	1.19	[0.69]	1.55	[0.93]	-1.77	[-2.64]	0.07	[0.08]	-1.50	[-2.15]
Current account	0.07	[0.40]	-0.24	[-1.26]	0.19	[1.03]	-0.64	[-1.18]	0.38	[0.71]	-0.14	[-0.25]	-0.58	[-2.51]	-0.16	[-0.66]	-0.55	[-2.40]
Short-term interest rate	0.06	[0.15]	0.75	[1.91]	-0.24	[-0.63]	-0.18	[0.51]	0.08	[0.16]	0.41	[0.79]	-0.69	[-1.79]	0.21	[0.58]	0.13	[0.35]

Notes: The values represent the coefficient (b) and the t-statistic of the shock variable in the following equation: $\Delta x(t) = cte + a1 * \Delta x(t\text{-}1) + a2 * \Delta x(t\text{-}2) + b * Shock(t)$. Shocks of the model with the short-run restrictions for the identification of public spending shocks and with inflation as the dependent variable.

Table 2

Correlation of structural shocks identified with different models

Belgium				**Ireland**				**Denmark**			
Supply shock				Supply shock				Supply shock			
	(p, l)	(rer, s)	(rer, l)		(p, l)	(rer, s)	(rer, l)		(p, l)	(rer, s)	(rer, l)
(p, s)	0.99	0.68	0.66	(p, s)	0.90	0.64	0.62	(p, s)	0.99	0.87	0.84
(p, l)		0.61	0.61	(p, l)		0.56	0.68	(p, l)		0.86	0.85
(rer, s)			0.99	(rer, s)			0.95	(rer, s)			0.93
Demand shock				**Demand shock**				**Demand shock**			
	(p, l)	(rer, s)	(rer, l)		(p, l)	(rer, s)	(rer, l)		(p, l)	(rer, s)	(rer, l)
(p, s)	0.95	0.48	0.51	(p, s)	1.00	0.51	0.51	(p, s)	0.99	0.50	0.48
(p, l)		0.48	0.54	(p, l)		0.48	0.48	(p, l)		0.48	0.51
(rer, s)			0.99	(rer, s)			1.00	(rer, s)			0.91
Public spending shock				**Public spending shock**				**Public spending shock**			
	(p, l)	(rer, s)	(rer, l)		(p, l)	(rer, s)	(rer, l)		(p, l)	(rer, s)	(rer, l)
(p, s)	0.95	0.97	0.96	(p, s)	0.89	0.97	0.92	(p, s)	0.99	0.98	0.96
(p, l)		0.93	0.96	(p, l)		0.90	0.94	(p, l)		0.98	0.97
(rer, s)			0.99	(rer, s)			0.95	(rer, s)			0.98

Notes: (p, s) represents the model with short-term restriction to identify the public spending cuts, and with the inflation variable. (rer, s) represents the model with short-term restriction to identify the public spending cuts, and with the changes in the real exchange rate as dependent variable instead of inflation. (p, l) represents the model with long-term restriction to identify the public spending cuts, and with the inflation variable. (rer, l) represents the model with long-term restriction to identify the public spending cuts, and with the change in the real exchange rate.

The resulting series for the shocks of the different model specifications should be related to illustrate their independence from the identification restrictions and the variables selected (Table 2). The government spending shocks are always very strongly correlated in both the short and long-run restrictions, and in both versions with inflation or with changes in real exchange rates. Supply and demand shocks are independent from the identification of the public spending shock. They differ, however, if inflation is replaced by real exchange rate changes. But the series for the supply shock are still highly correlated and as the demand and monetary shocks have different effects on inflation and the change in the real exchange rate, the overall result is acceptable.

The impulse response graphs show the macroeconomic reactions to the three types of shock (Figure 11). The reaction of growth and inflation to public spending shocks is of particular interest in this paper.

For Belgium, spending cuts do not have a significant effect on growth, nor on inflation. Gradually the reponse becomes negative, but only the negative effect on the price level is important in the long run. The real exchange rate does not show any strong reaction. So the nominal exchange rate appreciation more or less follows the price decrease, but this process develops slowly over time.

In Ireland, there is a strong and significant (for the first year only) negative impact on GDP following spending cuts, implying a real impact multiplier greater than one. Prices, on the other hand, do not react on impact but gradually decrease afterwards. The real exchange rate depreciates less than prices, so that the nominal exchange rate appreciates slightly.

In Denmark, too, the effect on output is strong and significant, but that on prices is small. As the government spending shock is followed by further cuts, output reacts relatively less in the long run, with very strong price effects. The real exchange rate does not show any movement in the long run, implying a strong appreciation following price declines (the absence of a relaxation in policy can explain the strong real multiplier and the strong price declines in this case).

Together, these results show that fiscal shocks have insignificant effects on prices in the short run. In fact, in all three countries prices increase on impact, but not significantly so. Strong price rigidities can explain both the small price reaction and the strong output effects in the short run. In Belgium, the reduction in interest rate differentials with Germany, following fiscal policy adjustments (improving the Belgian fundamentals), may explain the small output effects. In the long run, output and prices are affected negatively in all three countries. So the negative demand effect is not offset by a strong private demand response, implying that horizons were less than infinite. The theoretical hypothesis of the negative wealth effect on the labour supply is not rejected by the results.

The evidence from the impulse response analysis should be complemented with the forecast error variance decomposition (Table 3). These results illustrate the importance of the public spending shocks in explaining growth and inflation on average over the estimation period. In Belgium, public spending shocks make almost no contribution to the variance decomposition of growth or inflation. In Ireland, spending shocks explain some 20% of the variance of growth but they are not important for inflation (with the long-run restriction to identify spending shocks, fiscal shocks become much more important for growth). In Denmark, spending shocks are even more important and explain more than 30% of the variance of growth, and they are the dominant source of disturbances in inflation in the long run.

For Belgium, Ireland and Denmark we also show the contribution of the different shocks to output growth over the period 1980-96 (Figure 12). This should indicate the relative role of the different shocks in explaining economic performance over this period, and especially during the fiscal adjustment programmes. With these results, we are able to describe the role of the public spending cuts during the adjustment process in the 1980s.

When the adjustment programme started in Ireland in 1987, output was very low because of unfavourable supply conditions during the first half of the 1980s. The tax increases during that period do not appear in the demand shocks, but probably worked through the negative supply shocks. The expenditure cuts that were undertaken in 1987-89 are very evident in the spending shocks. They

Figure 11a

Impulse response of the SVAR model with inflation and a short-run restriction for the identification of public expenditure shocks

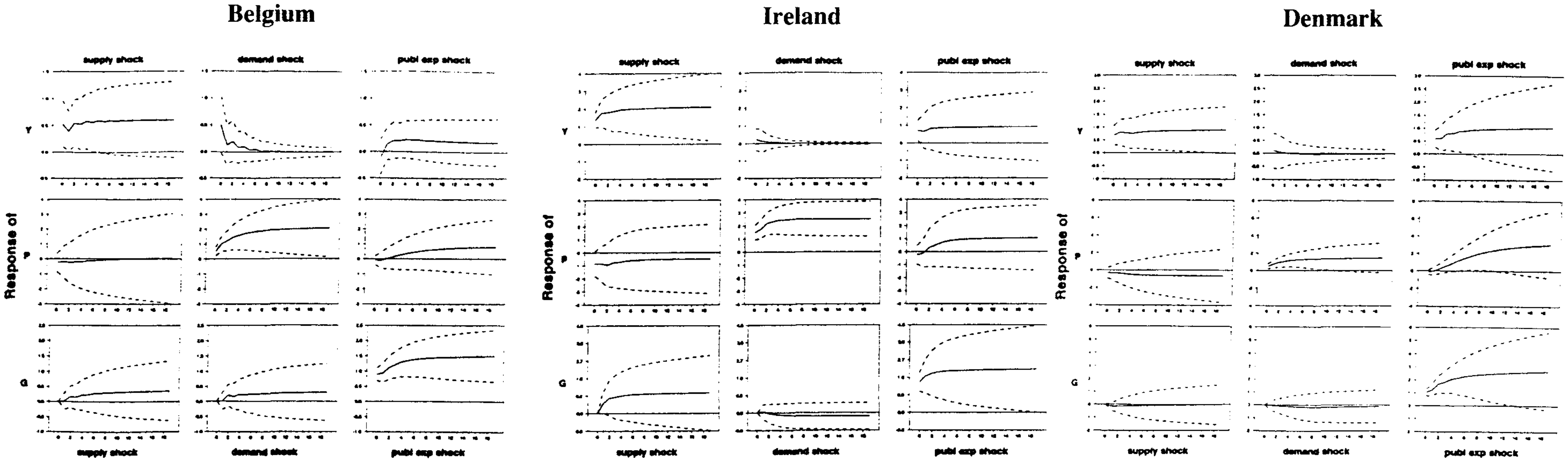


Figure 11b

Impulse response of the SVAR model with real exchange rate changes and a short-run restriction for the identification of public expenditure shocks

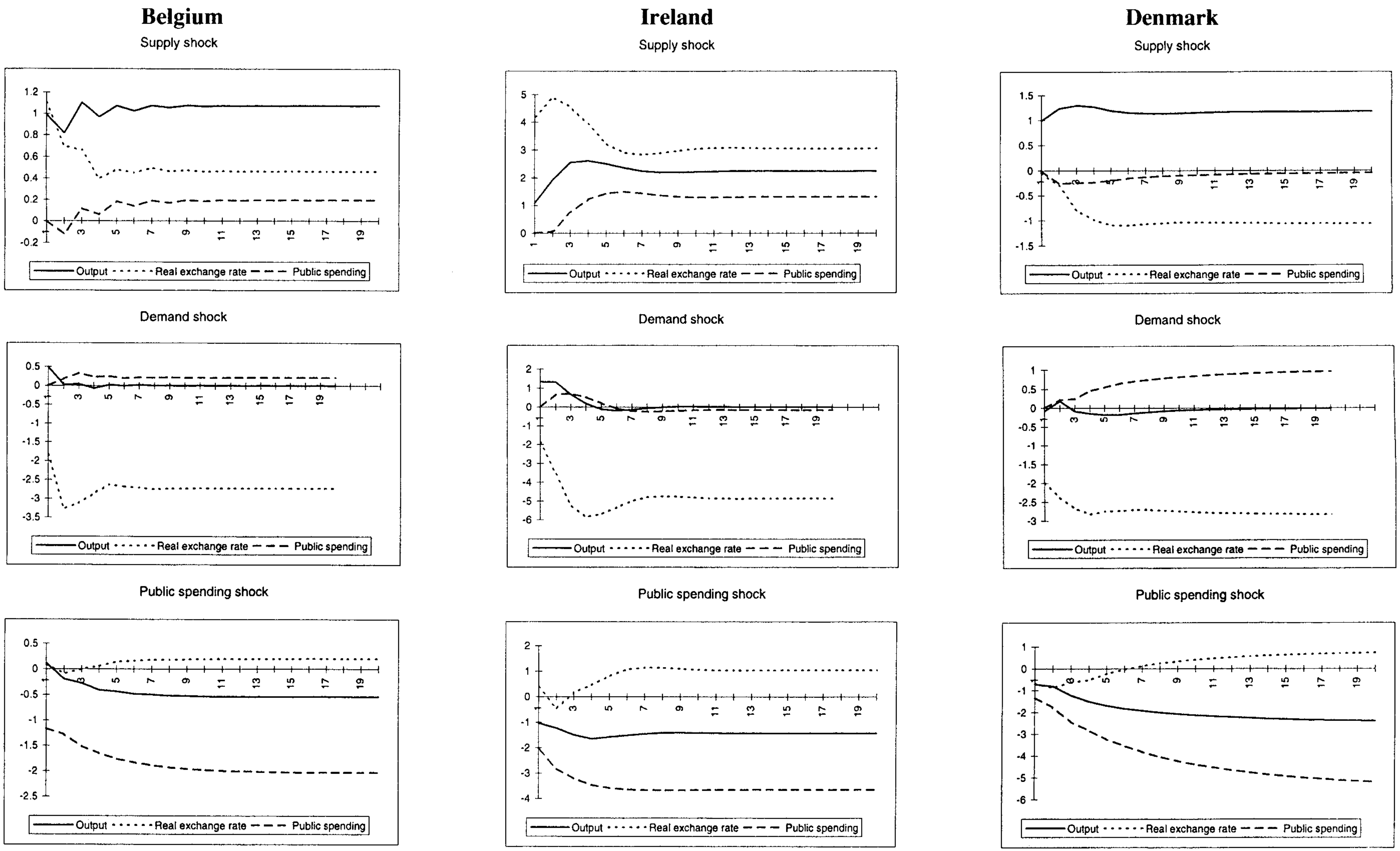
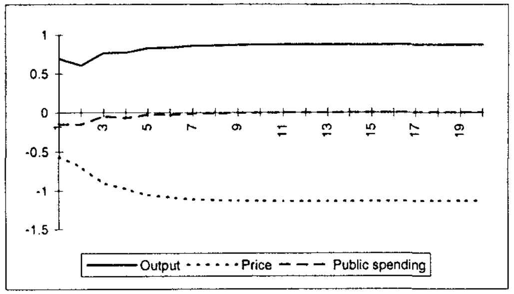


Figure 11c

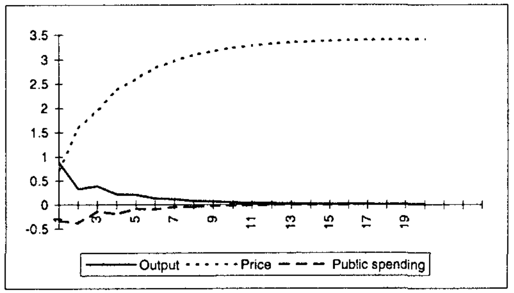
Impulse response of the SVAR model with inflation and a long-run restriction for the identification of public expenditure shocks

Belgium

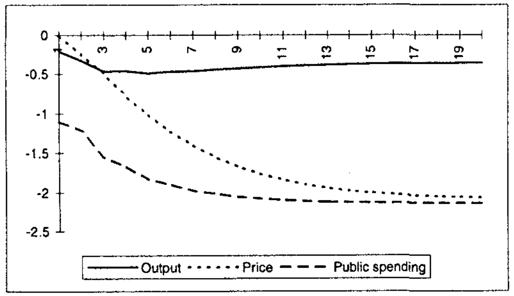
Supply shock



Demand shock

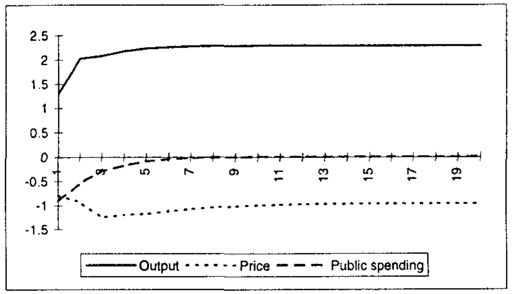


Public spending shock

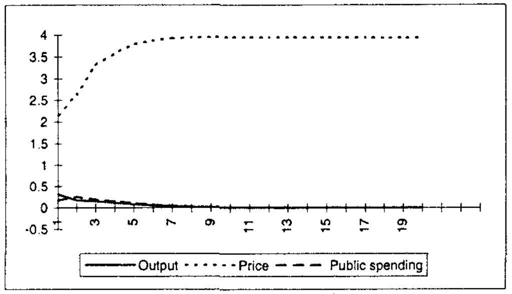


Ireland

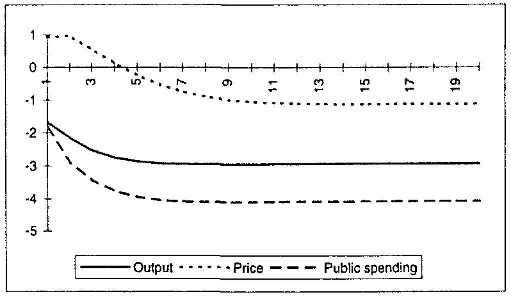
Supply shock



Demand shock

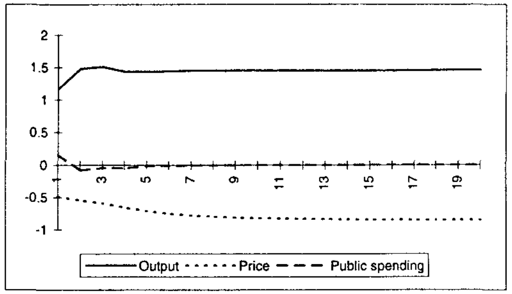


Public spending shock

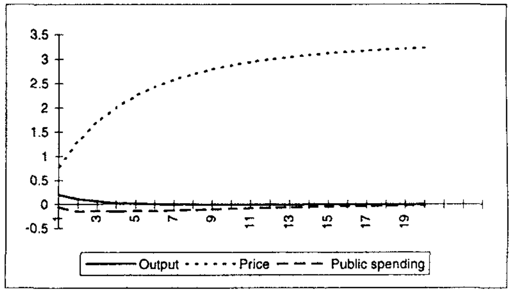


Denmark

Supply shock



Demand shock



Public spending shock

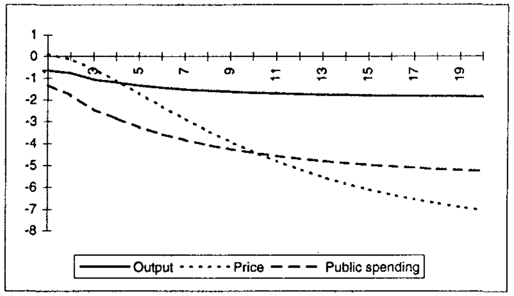
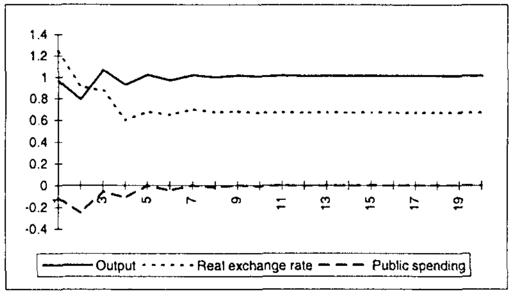


Figure 11d

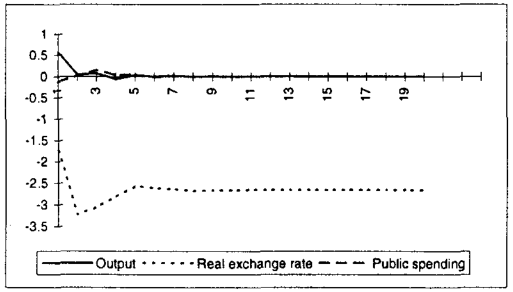
Impulse response of the SVAR model with real exchange rate changes and a long-run restriction for the identification of public expenditure shocks

Belgium

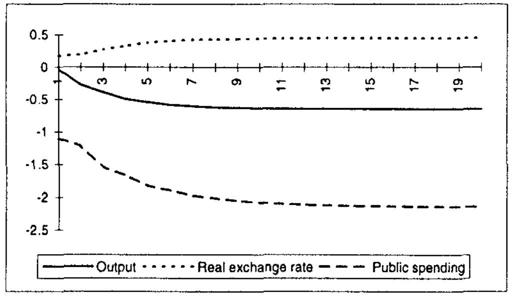
Supply shock



Demand shock

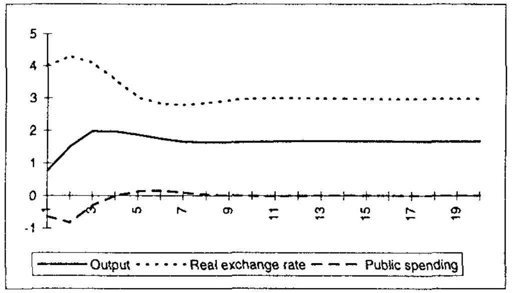


Public spending shock

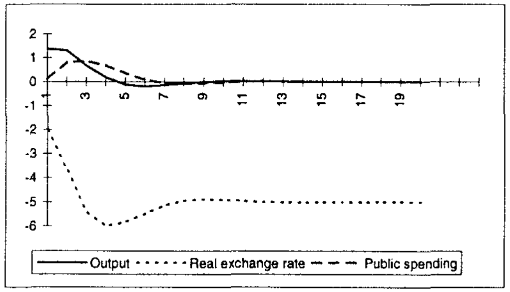


Ireland

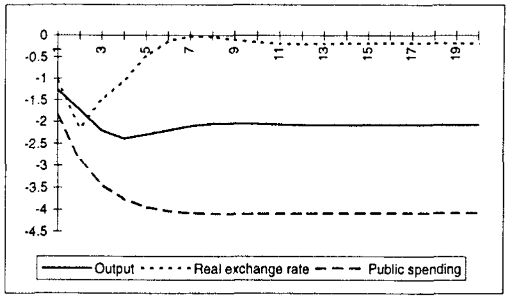
Supply shock



Demand shock

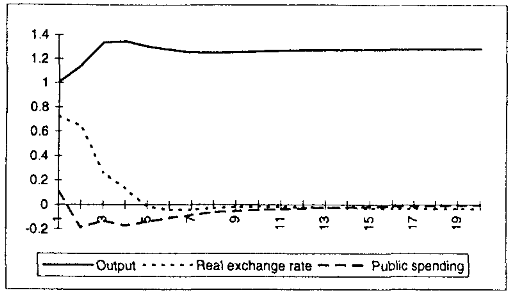


Public spending shock

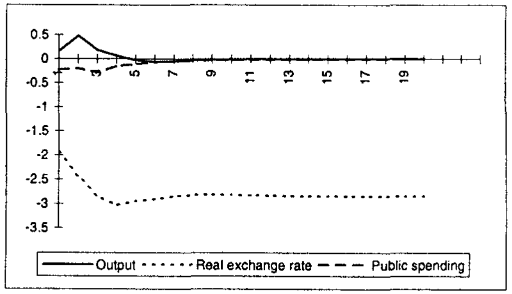


Denmark

Supply shock



Demand shock



Public spending shock

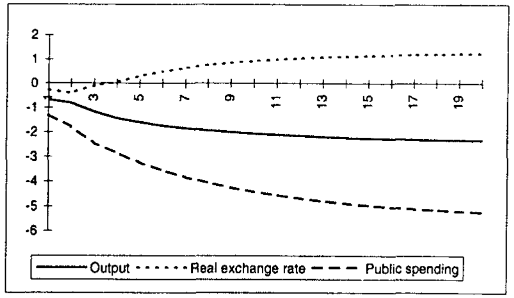


Table 3

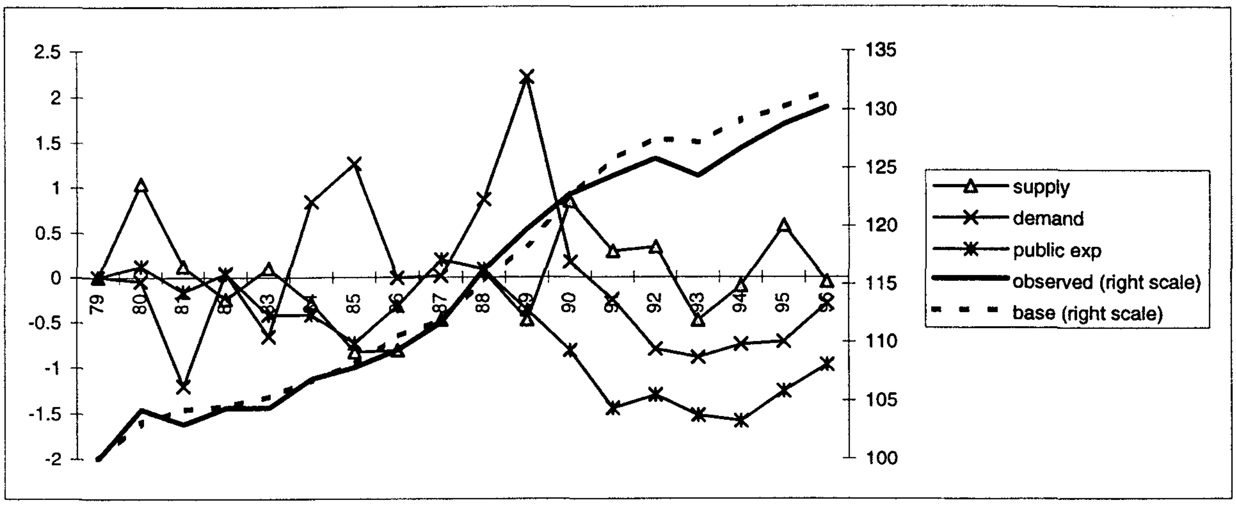
Variance decomposition of the forecast error: model with inflation and short-run restrictions

Belgium				Ireland				Denmark			
GDP growth				GDP growth				GDP growth			
Years	**Supply**	**Demand**	**Public exp.**	**Years**	**Supply**	**Demand**	**Public exp.**	**Years**	**Supply**	**Demand**	**Public exp.**
1	46.58	51.82	1.59	1	77.93	2.07	20.00	1	65.11	1.05	33.84
2	37.85	55.86	6.29	2	80.69	2.02	17.29	2	65.49	1.59	32.92
3	38.74	54.65	6.61	3	79.41	1.99	18.61	3	62.42	1.56	36.02
4	38.07	55.33	6.60	4	79.16	2.01	18.83	4	62.18	1.57	36.24
5	38.18	55.20	6.62	5	79.10	2.02	18.88	5	61.53	1.56	36.91
6	38.04	55.37	6.60	6	79.08	2.03	18.88	6	61.32	1.55	37.13
7	38.04	55.36	6.59	7	79.08	2.04	18.88	7	61.11	1.55	37.34
8	38.00	55.41	6.59	8	79.08	2.04	18.88	8	61.01	1.54	37.44
9	37.99	55.41	6.59	9	79.08	2.04	18.88	9	60.94	1.54	37.52
10	37.98	55.42	6.60	10	79.08	2.04	18.88	10	60.89	1.54	37.57
Inflation				Inflation				Inflation			
Years	**Supply**	**Demand**	**Public exp.**	**Years**	**Supply**	**Demand**	**Public exp.**	**Years**	**Supply**	**Demand**	**Public exp.**
1	29.69	68.33	1.98	1	23.54	74.93	1.53	1	18.96	77.01	4.03
2	14.93	84.07	1.00	2	23.03	75.43	1.54	2	13.67	80.16	6.17
3	14.17	83.75	2.08	3	20.88	73.23	5.89	3	10.77	73.53	15.70
4	12.49	84.47	3.04	4	20.85	71.62	7.53	4	9.19	64.93	25.88
5	11.87	84.12	4.01	5	20.73	70.31	8.95	5	8.08	57.04	34.89
6	11.40	83.82	4.78	6	20.87	69.50	9.62	6	7.25	50.87	41.88
7	11.17	83.48	5.35	7	20.96	69.08	9.96	7	6.63	46.17	47.20
8	11.02	83.22	5.75	8	21.02	68.88	10.10	8	6.17	42.66	51.17
9	10.94	83.04	6.02	9	21.05	68.79	10.16	9	5.83	40.03	54.14
10	10.89	82.91	6.20	10	21.07	68.76	10.18	10	5.58	38.06	56.36
Public expenditure				Public expenditure				Public expenditure			
Years	**Supply**	**Demand**	**Public exp.**	**Years**	**Supply**	**Demand**	**Public exp.**	**Years**	**Supply**	**Demand**	**Public exp.**
1	0.00	0.00	100.00	1	0.00	0.00	100.00	1	0.00	0.00	100.00
2	0.01	0.00	99.99	2	11.10	0.05	88.84	2	4.30	0.09	95.61
3	1.93	6.25	91.82	3	14.28	0.16	85.56	3	3.58	0.18	96.24
4	1.92	6.19	91.89	4	15.01	0.19	84.80	4	3.46	0.18	96.36
5	2.24	7.24	90.52	5	15.32	0.23	84.45	5	3.26	0.21	96.54
6	2.23	7.22	90.54	6	15.40	0.25	84.35	6	3.19	0.22	96.59
7	2.29	7.43	90.28	7	15.42	0.27	84.31	7	3.13	0.24	96.63
8	2.29	7.43	90.28	8	15.43	0.27	84.30	8	3.10	0.25	96.65
9	2.30	7.47	90.22	9	15.43	0.27	84.30	9	3.07	0.26	96.67
10	2.30	7.48	90.22	10	15.43	0.27	84.30	10	3.05	0.28	96.67

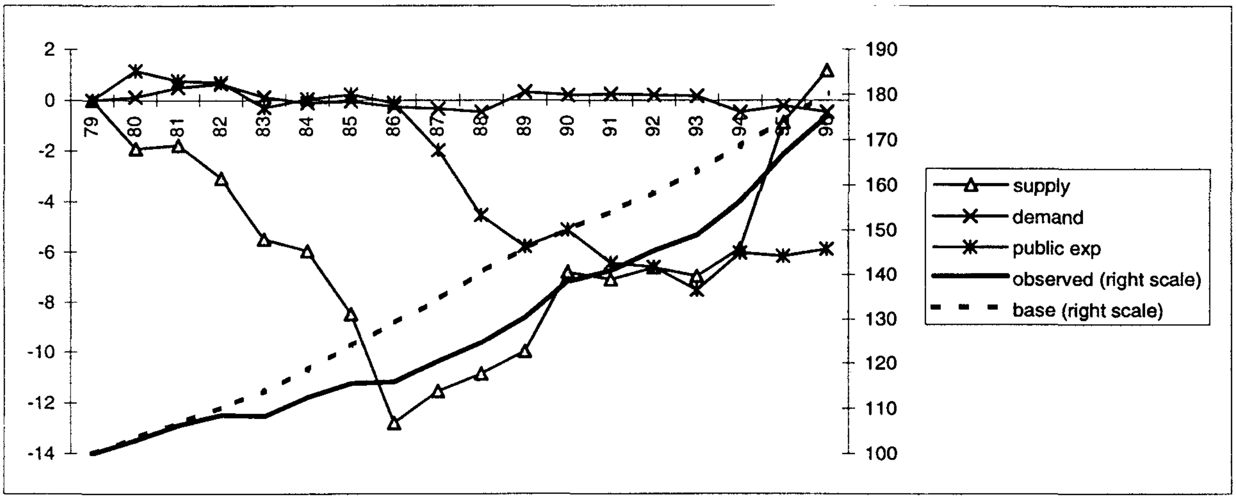
Figure 12

GDP evolution over the period 1980–96:
contribution of the three shocks in the explanation of gap to observed – base

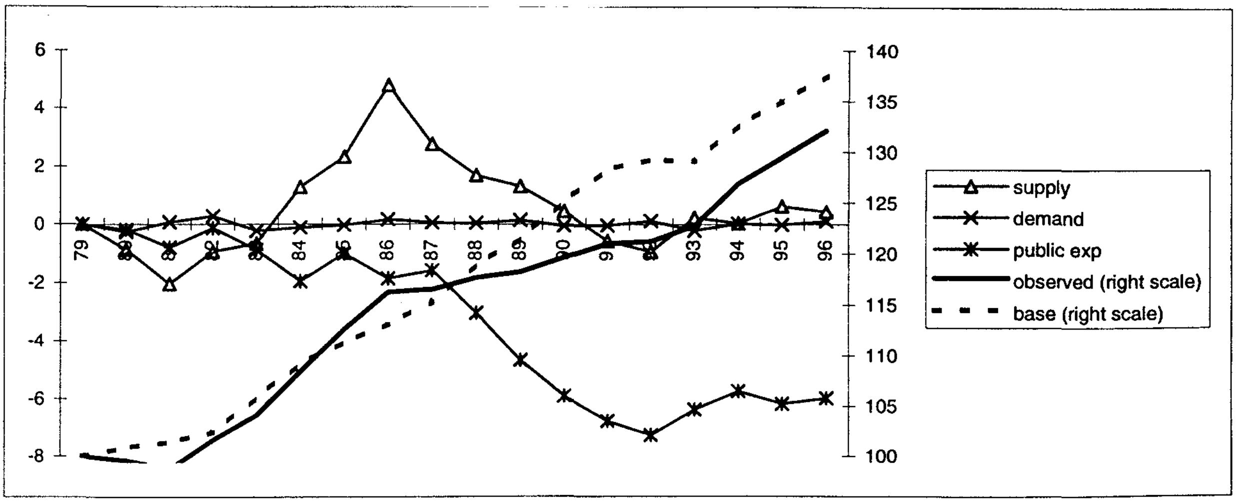
Belgium



Ireland



Denmark



had a very strong negative effect on economic growth, and that effect remained present until the end of our estimation period (1996). In the model with inflation, the negative influence of public spending cuts is neutralised by major positive supply shocks that not only compensated for the negative influence of public spending but also allowed the gap that was built up at the beginning of the 1980s to be closed. However, in the model with the real exchange rate, there were major positive demand shocks from 1986 to 1990. Supply shocks only occurred in the 1990s. These results therefore point to the importance of the Irish depreciation in 1986 in offsetting the negative public spending cuts. These results contradict the hypothesis of Giavazzi and Pagano, in which it was positive domestic private demand shocks, following the positive wealth effect, that were responsible for the overall positive outcome of the stabilisation programme. Our results are in accordance with the remarks of Barry and Devereux, who claim that the Irish success was due to shocks other than those in public expenditure.

Denmark experienced substantial spending cuts in the periods 1983-84 and 1988-91. During the first period, the shocks were offset by positive supply innovations. During the second, there were no offsetting shocks and growth remained below its normal growth path. The fiscal shocks contributed to the good inflation record in Denmark.

In Belgium, public spending shocks occurred in 1982 and in 1987-90, according to the model. The negative influence on GDP was relatively small. The impact on inflation was greater and, as in the Danish case, it contributed to the good performance in terms of inflation in the 1990s.

Conclusion

General equilibrium models offer a suitable framework for analysing the impact of fiscal consolidation programmes for small open economies. Different arguments that are encountered in the literature and in empirical macroeconomic model simulations can be reproduced with these theoretical coherent models. Simulation exercises allow us to indicate more precisely the specific assumptions behind some results such as the "expansionary fiscal contraction". These exercises also reveal the importance of supply conditions, monetary policy reactions and exchange rate behaviour in determining the outcome of fiscal shocks, especially in the context of small open economies with price rigidity in the short run.

Although the empirical significance of the SVAR results is low, there is some evidence that government expenditure cuts had short-term negative demand effects on output in countries such as Belgium, Denmark and Ireland. This result contradicts the hypothesis of large positive wealth effects following the fiscal contractions in these countries. Our decomposition provides some support for the hypothesis that simultaneously there were positive supply shocks at work that offset the negative demand effects and were responsible for the overall positive growth effects.

Further research should be oriented towards a better integration of the theoretical model and the empirical evidence. Therefore, a fully calibrated general equilibrium model is needed. Within the theoretical model one should pay more attention to a realistic representation of the labour market and the monetary policy reaction function, as the interaction between public spending shocks and these behavioural functions is crucial for the outcome of the shocks. Especially in the context of small open economies, examination of these channels would seem to be more important than the further elaboration of specific wealth effects following fiscal consolidations.

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Comments on: "Fiscal consolidation in general equilibrium models" by Raf Wouters

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by Filippo Altissimo

The goal of this work is to assess the effects of sizeable fiscal consolidation in a small open economy, focusing in particular on the recent experience of the Belgian economy.

The author starts by reviewing the large body of literature on fiscal consolidation and the effects of government spending cuts. The survey begins by looking at the effects of fiscal consolidation in a simple static Keynesian model; more general dynamic general equilibrium models are then tackled. According to both approaches, the effects of fiscal consolidation are mixed in terms of intensity; in neither of them, however, a fiscal consolidation has expansive effects. On the other hand, recent studies by Alesina and Perotti (1996), Blanchard (1990), Bertola and Drazen (1993) and others suggest that fiscal consolidation can have expansive effects. In these studies the positive feedback on the economy stems from the fact that fiscal cuts imply expectations of a lower future tax burden.

The author argues that, even if non-Keynesian mechanisms have indeed played a role in some fiscal consolidation experiences, they are difficult to disentangle and identify. In particular, the outcomes of the episodes of expansive fiscal consolidation which are usually cited in the literature (Ireland and Denmark), can be attributed to a combination of policies, whose individual contribution cannot be easily assessed. Thus the aim of the present study is to identify the policy mix which contributed to the success of the consolidations in Ireland and Denmark, and to compare those experiences with the Belgian one. The paper tackles the issue in two different ways: first, a general equilibrium model of consumers and firms is specified and various policy simulations are carried out; second, trivariate structural VARs for the Belgian, Irish and Danish economies are estimated.

The model proposed by the author is characterised by two types of agents: agents of the first kind are liquidity constrained, while the others are not; the agents solve an infinite horizon welfare maximisation problem; two different consumption goods are available. On the production side, firms act as monopolistic competitors and face a quadratic cost adjustment problem. The model is used to simulate the effects of different policy shocks. The exercise is performed assuming both transitory and permanent shocks, and under a variety of hypotheses concerning labour supply behaviour and price rigidity. The results of the exercise are in line with the expectation that the effect of fiscal consolidation on output is negative. The overall negative effect is, however, damped by an increase in private consumption. The size of the effect is a function of the assumptions regarding the parameter values. It is, however, unclear whether the policy simulations are meant to summarise the various results which can be found in the literature on fiscal consolidation, or whether their aim is to represent the behaviour of the Belgian economy specifically. This ambiguity is partly due to the fact that the author does not describe how the structural parameters of the model have been chosen and whether they have been calibrated to the Belgian data.

In the second part of the paper the author estimates structural VARs describing GNP, inflation and government expenditure, using a sample of 35 observations of annual data for the three countries of interest. The exact identification of the structural VARs requires three restrictions on the error structure of the VARs. The first one is the usual identification condition for demand shocks, which requires that demand shocks have no long run effect on output. The remaining two conditions are related to the effects of demand and supply shocks on public finance. The author experiments with two different specifications for the latter two identification restrictions: the first one imposes a minimum delay restriction on the effects of demand and supply shocks on public finance; the second one implies no long run effects of demand and supply shocks on public finance. A clear cut preference for one of the proposed identification schemes is however not provided; the resulting ambiguity makes it difficult to assess the underlying economic interpretation of the proposed identification schemes.

More generally, the way in which the identification of the VARs is assessed is not convincing. To this end, the work follows two different paths. First, the structural shocks are projected on other macro variables and identification is judged on the basis of the signs of the contemporaneous correlation. Second, the correlation of structural shocks across different identification specifications is examined. The following objections can be raised concerning these approaches to the identification issue. First, there is a very vague relation between the signs of the projections and the identification scheme. Second, this way of assessing does not consider the fact that the identification conditions imposed rely both on short and long run restrictions. Third, the economic interpretation should drive the choice of the identification scheme and not the other way.

Given the identification schemes, the VARs are estimated twice, first with inflation being included in the model, then with the real exchange rate replacing inflation. The reason for doing the latter is that the real exchange rate may be a better proxy of the monetary policy stance. However, if the final aim of the work is to disentangle the different sources of a successful fiscal consolidation, it would be more appropriate to work with a better articulated model, which should include at least an explicit measure of the monetary stance, for example as in the work of Bernanke and Mihov (1996).

To sum up, according to the cited literature on the non-Keynesian effect, the positive effect of fiscal consolidation can result either from the working of the expectation mechanism or from the presence of a trigger point in the decision of economic agents beyond which the postponement of fiscal action becomes counterproductive. Those non-linearities in the decision process of economic agents, which are needed for non-Keynesian effects to be possible, are unlike to be identified by means of VAR models. Obviously, the linearity of the VAR implies that the response to a negative fiscal shock must be the same independently of the history of the model at the time when the fiscal contraction occurs. The use of VARs as a means to analyse the policy mixes in the three countries, and in particular to highlight the importance of the non-Keynesian effects, may thus be inappropriate.

Can VARs describe monetary policy?

Charles L. Evans and Kenneth N. Kuttner*

Introduction

How does monetary policy affect the economy? To answer that question requires solving a basic simultaneity problem: monetary policy affects the economy while at the same time responding endogenously to changing macroeconomic conditions. Empirically estimating the effects of policy therefore requires some observable exogenous element to policy. The narrative approach pioneered by Friedman and Schwartz (1963), and applied by Romer and Romer (1989), is one way to identify exogenous policy shifts. A more common approach, however, is the Vector Autoregression (VAR) technique developed by Sims (1972, 1980a and 1980b). In this procedure, changes in the monetary policy instrument that are not explained by the variables included in the model are interpreted as exogenous changes in policy, or policy "shocks". Christiano, Eichenbaum and Evans (1997) provides a survey of this line of research.

One unresolved question is how well simple econometric procedures, like VARs, can describe the monetary authority's response to economic conditions, and by extension, the policy shocks used to identify policy's effects on the economy. There are several reasons to be skeptical of the VAR approach. VAR models (indeed, all econometric models) typically include a relatively small number of variables, while the Fed is presumed to "look at everything" in formulating monetary policy. By assuming linearity, VARs rule out plausible asymmetries in the response of policy, such as those resulting from an "opportunistic" disinflation policy. VAR coefficients are assumed to remain constant over time, despite well-documented changes in the Fed's objectives and operating procedures.

The goal of this paper is to assess VAR accuracy in predicting changes in monetary policy, and the shock measures derived from those predictions, using forecasts from the Federal funds futures market as a basis for comparison. Section 1 reviews the VAR methodology, and its putative deficiencies. Section 2 shows that the correlation between the VAR and the futures market forecast errors can give a misleading picture of the VAR's accuracy, and suggests looking instead at the relationship between the forecasts themselves. The results in Section 3 show that profligacy *per se* does not hurt the VAR's performance, but reducing the number of lags *and* estimating the model over a more recent subsample can improve the model's forecast accuracy. Section 4 discusses the time aggregation problems inherent in extracting policy surprises from Fed funds futures data. The conclusions, summarized in the final section, are that VARs mimic the futures market's forecasts reasonably well, but that it would be misleading to use futures market forecasts as the only basis for comparison.

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1. VARs: the technique and its critiques

The reduced form of a VAR simply involves the regression of some vector of variables, x, on lags of x:

$$x_t = A(L))\, x_{t-1} + u_t \tag{1}$$

$A(L)$ is a polynomial in the lag operator, and u_t is a vector of disturbances, where $E(u_t\, u'_t) = \Omega$. In monetary applications, the x_t vector would include one (or more) indicators of monetary policy, along with the other macroeconomic and financial variables. The widely-used model of Christiano, Eichenbaum and Evans (CEE) (1996a), includes the Federal funds rate, r_t, logarithms of lagged payroll employment (N), the personal consumption deflator (P), nonborrowed reserves (NBR), total reserves (TR), $M1$, and the smoothed growth rate of sensitive materials prices ($PCOM$).[1] The funds rate equation from the CEE model is:

$$r_t = \beta_0 + \sum_{i=1}^{12} \beta_{1,i} r_{t-i} + \sum_{i=1}^{12} \beta_{2,i} \ln N_{t-i} + \sum_{i=1}^{12} \beta_{3,i} \ln P_{t-i} + \sum_{i=1}^{12} \beta_{4,i} PCOM_{t-i}$$

$$+ \sum_{i=1}^{12} \beta_{5,i} \ln NBR_{t-i} + \sum_{i=1}^{12} \beta_{6,i} \ln TR_{t-i} + \sum_{i=1}^{12} \beta_{7,i} \ln M1_{t-i} + u_t$$

a regression of the Fed funds rate on lagged values of all the variables included in the VAR.

In the "structural" VAR:

$$x_t = B_0\, x_t + B(L)\, x_{t-1} + e_t \tag{2}$$

the shocks and feedbacks are given economic interpretations. The covariance matrix of the e innovations is diagonal, and contemporaneous feedback between elements of x is captured by the B_0 matrix. The equation involving the monetary policy instrument (the Fed funds rate, for example) is often interpreted as a reaction function describing the Fed's response to economic conditions. By extension, the innovation to this equation is taken to represent "shocks" to monetary policy. A great deal of research and debate has centered on the identifying assumptions embodied in the choice of B_0. Examples include Bernanke (1986), Sims (1992), Strongin (1995), Bernanke and Mihov (1995), Christiano, Eichenbaum and Evans (1996b), and Leeper, Sims and Zha (1996). The typical focus of this research is the impulse response functions of macroeconomic variables in response to monetary policy shocks, and how the specific identifying assumptions affect the responses.

This paper does *not* deal with the identification issue, but focuses instead on a more basic question: whether the reduced form of the VAR, and the monetary policy equation in particular, generate sensible forecasts. Rudebusch (1997) pointed out that the Fed funds futures market provides a ready benchmark for evaluating the VAR's Fed funds rate equation.[2] This strategy makes sense if the futures market is efficient, in the sense that its errors are unforcastable on the basis of available information.[3] The month $t-1$ one-month futures rate, $f_{1,t-1}$, would therefore represent the rational

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1 This series was a component in the BEA's index of leading economic indicators prior to its recent revision by the Conference Board. Recent observations of the series used here are computed using the BEA's methodology from raw commodity price data from the Conference Board.

2 Fed funds futures are known officially as Thirty-Day Interest Rate futures.

3 The findings of Krueger and Kuttner (1996) generally support this view. Results less supportive of market efficiency were reported by Sims (1996). In a regression of the average Fed funds rate on lagged monthly averages of T-bill and discount rates, and Fed funds futures rate from the middle of the previous month, the T-bill and discount rates were

Table 1

The case against VARs

<table>
<tr><th rowspan="3">Horizon, in months</th><th colspan="4">Standard deviation</th><th rowspan="3">Average standard error of CEE shock</th><th rowspan="3">CEE-futures correlation</th></tr>
<tr><th rowspan="2">Futures</th><th rowspan="2">No change</th><th colspan="2">CEE model</th></tr>
<tr><th>in sample</th><th>out of sample</th></tr>
<tr><td>One</td><td>13.1</td><td>19.2</td><td>23.5</td><td>27.1</td><td>20.6</td><td>0.35</td></tr>
<tr><td>Two</td><td>20.8</td><td>33.1</td><td>35.4</td><td>43.8</td><td>34.0</td><td>0.38</td></tr>
<tr><td>Three</td><td>39.6</td><td>46.8</td><td>48.2</td><td>60.7</td><td>42.1</td><td>0.47</td></tr>
</table>

Notes: The CEE model is estimated on monthly data from January 1961 through July 1997. The reported statistics are based on the May 1989 through December 1997 sample. The in-sample standard deviation and correlations are adjusted for the degrees of freedom used in estimation. Units are basis points.

expectation of the period t Fed funds rate conditional on information at time t–1. The corresponding surprise would be:

$$u^*_t = r_t - f_{1,t-1}$$

where r_t is the *average* overnight Fed funds rate in month t, consistent with the structure of the futures contract. To avoid familiar time averaging problems, the futures rate is taken from the last business day of the month.[4]

A glance at the series plotted in Figure 1 shows that the VAR's forecast errors bear little resemblance to the futures market surprises. The correlation between the two is only 0.35 for one-month-ahead forecasts, comparable to the R^2 of 0.10 reported in Rudebusch (1997).[5] The correlations between two- and three-month ahead shocks are somewhat higher.[6] If the futures market surprises are interpreted as the "true" shocks, this immediately calls the VAR approach into question.

A related problem is the VAR's poor forecasting performance, both in and out of sample. As shown in Table 1, the standard deviation of the regression's residuals is much higher than the futures market's forecast errors. The out-of-sample RMSE is higher still, well in excess of the RMSE of a naive "no change" forecast.

A third, less widely recognized, problem is the large standard error associated with the VAR's policy shocks. The variance of the estimated shocks, $\hat{u}$ around the "true" errors, u, is simply:

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statistically significant. When comparable point-sampled interest rates are used as regressors, however, market efficiency (i.e., the joint hypothesis that the coefficient on the lagged futures rate is 1 and the coefficients on other interest rates are zero) cannot be rejected.

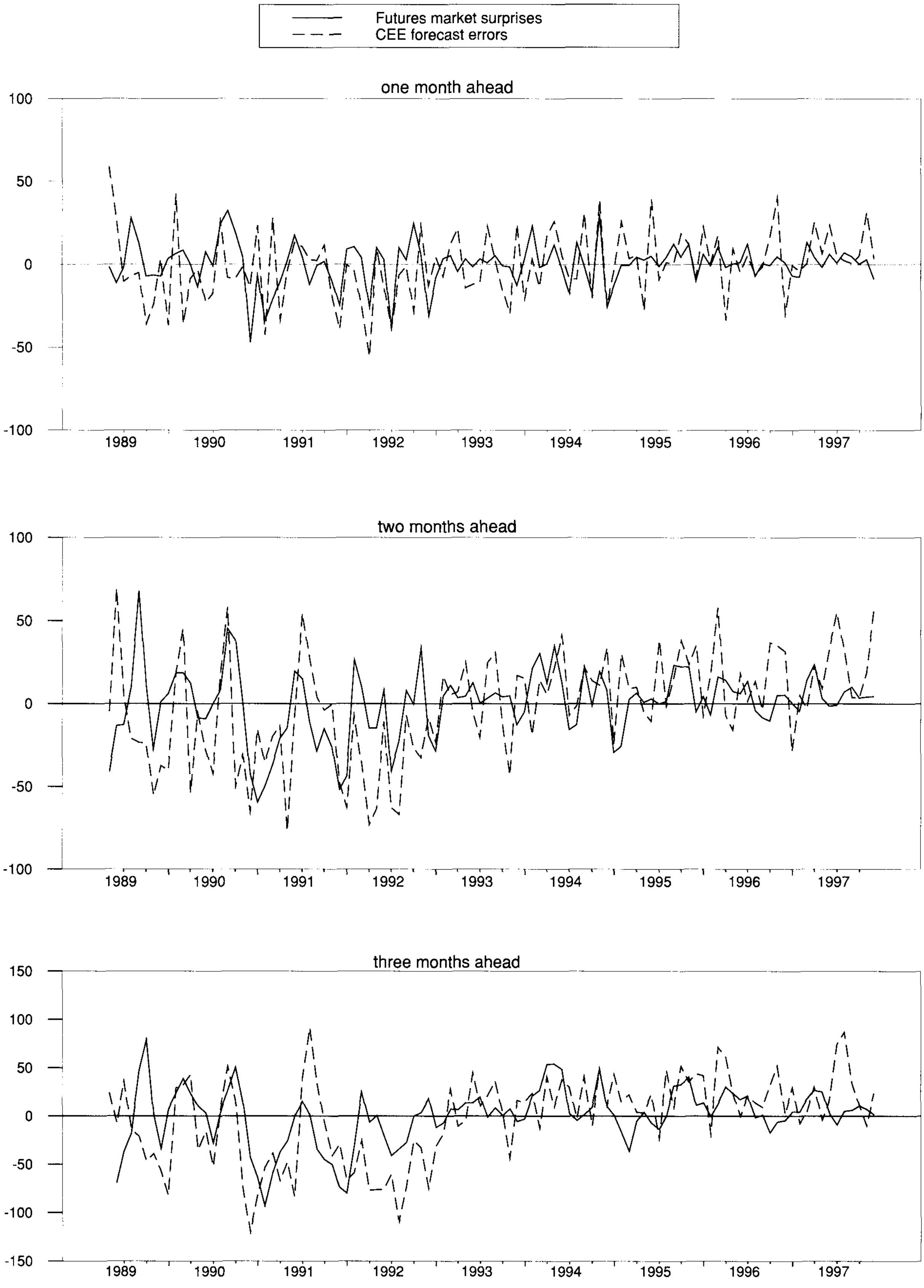
4 The futures rate appears to contain a forward premium of approximately 5 basis points for the one-month contract, which is subtracted from the futures rate in calculating the forecast error.

5 The variance and covariance estimates used to compute correlation coefficients are adjusted for the degrees of freedom used in estimation. Because computing the variance of the Fed funds futures shock does not require estimating any parameters, the result is equivalent to multiplying the unadjusted correlation by the factor $\sqrt{T/(T-k)}$ where T is the number of observations, and k is the number of VAR coefficients.

6 Two- and three-month ahead forecasts are obtained by regressing r_t on lags 2–12 and 3–12 of the same set of right-hand-side variables. The advantage of this shortcut is that it does not require estimating the entire VAR.

Figure 1

Fed funds futures surprises and CEE forecast errors

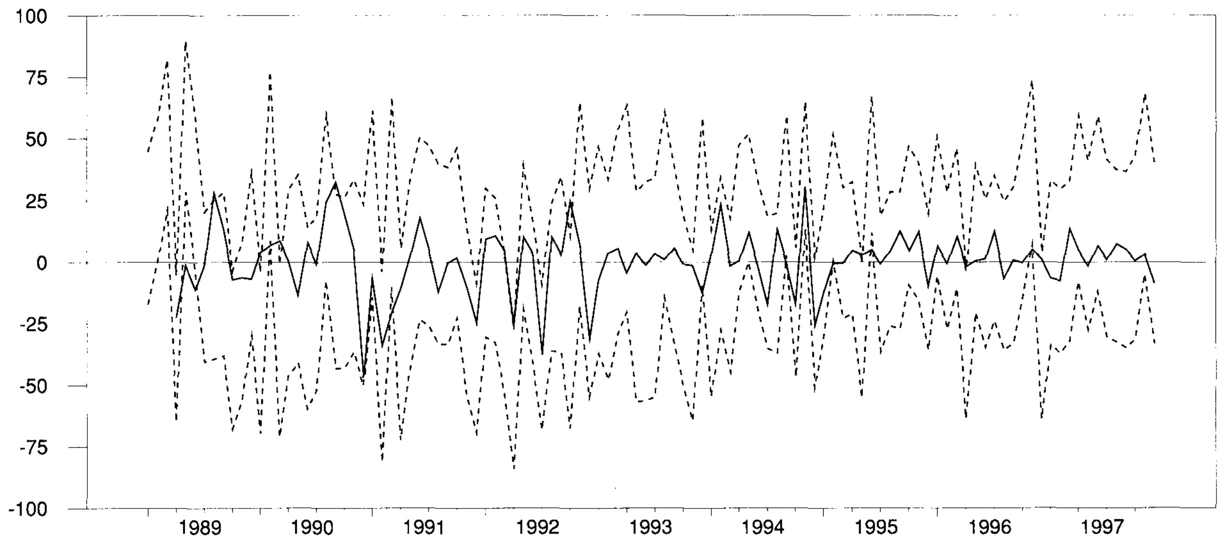


$$\begin{aligned}\mathrm{Var}(\hat{u}-u) &= E[(\hat{u}-u)(\hat{u}-u)'] \\ &= E\left[X(X'X)^{-1}X'uu'X(X'X)^{-1}X'\right] \\ &= \sigma^2 X(X'X)^{-1}X'\end{aligned}$$

where X is the $T \times k$ matrix of variables appearing on the right-hand side of the Federal funds rate equation, and σ^2 is the variance of u.[7] The average standard error of the shocks, reported in the fifth column of Table 1, and 95% confidence bounds around the CEE shocks are plotted in Figure 2. The estimates' imprecision is immediately visible in the figure; zero is well within the confidence bounds for most of the shocks, as are most of the futures market surprises.[8]

Figure 2

One-month Fed funds futures surprises and 95% CEE confidence bands



What accounts for this these deficiencies? Several authors, notably Rudebusch (1997), Pagan and Robertson (1995), and McNees (1992), have emphasized parameter instability as a possible explanation. Rudebusch cites the VAR's profligate specification as another candidate. Nonlinearities in the Fed's reaction function, such as those arising from "opportunistic" disinflation à la Orphanides and Wilcox (1996) are another possibility, and McCarthy (1995) provides some evidence supporting this view. In addition, the VAR leaves out variables that might help forecast monetary policy. Alternatively, the estimated equation may include variables not available to investors in real time.

One response to these criticisms is the conjecture that they do not matter for impulse response functions and variance decompositions, which are the focus of most VAR analyses. After all, VARs deliver robust, relatively precise estimates of the impact of monetary policy shocks while explicitly accounting for the shock estimates' uncertainty in the computation of the impulse response functions' error bands. Unfortunately, Fed funds futures rate data do not go back far enough to make reliable comparisons between impulse response functions based on VARs with those derived from futures market shocks. Christiano, Eichenbaum, and Evans (1997) report responses based on futures

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7 Since the regressors are not exogenous, this represents the variance of the posterior distribution conditional on realized X, given a flat prior.

8 The share of futures market surprises falling outside the bounds is 0.17, which represents a statistically significant deviation from the expected 0.05.

rate shocks, but with less than nine years of futures rate data, the standard errors are very large. Related work by Brunner (1997), however, found that shocks incorporating financial market and survey expectations generate impulse response functions similar to those from VARs, despite a low correlation between the measures.

Our view is that the problems are less severe than they appear, and that standard specifications can provide a better description of monetary policy than Rudebusch's results would suggest. First, the correlation between VAR and futures-market shocks is a poor gauge of the VAR's performance. Quantitatively small deviations from perfect futures market efficiency create a significant downward bias in the correlation. Second, small modifications to standard VAR specifications, such as reducing lag lengths and estimating over shorter samples, can tangibly improve the fit and precision of the models' forecasts. Finally, a time aggregation problem inherent in the futures rate can distort the timing and magnitude of shocks derived from the futures market.

2. How sensible is the correlation metric?

In using the Fed funds futures rate to evaluate the VAR's performance, one natural comparison is between the forecasts themselves: in this case, between the lagged one-month-ahead futures rate $f_{1,t-1}$ and the VAR's forecast $\hat{r}_t^{VAR}$. Perhaps because of the recent emphasis on policy shocks, however, assessments of VAR's performance have often involved the forecast errors rather than the forecasts themselves; see, for example, Rudebusch (1997), Brunner (1997) and Christiano et al. (1997).

At first glance, the correlation between shocks would seem to be a sensible basis for comparison; if the two procedures yielded the *same* forecasts, the shocks would be identical, and the correlation would be 1.0. Closer scrutiny shows that this measure can give a misleading picture, however; the covariance between the *shocks* has little to do with the covariance between the *forecasts*. The correlation between shocks can therefore make bad forecasting models look good, and good models look bad.

Table 2

Alternative measures of fit

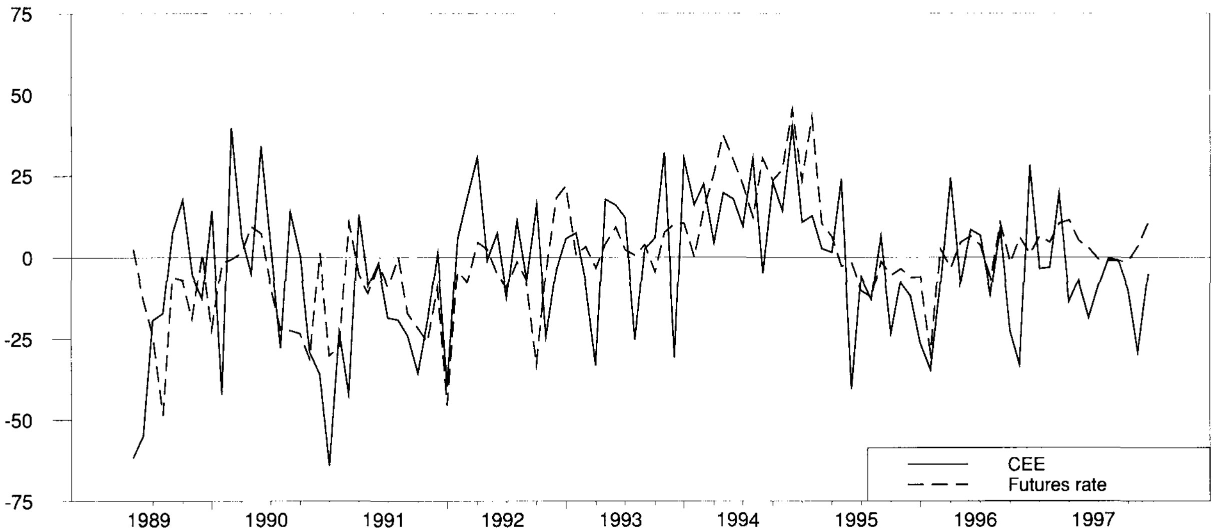
Model	Shock correlation	Δr forecasts	
		Correlation	Regression $\hat{b}$
CEE	0.35	0.43	0.57
No change	0.52	0	0

A comparison between the VAR and a naive "no change" forecast forcefully illustrates this point. Because the Fed funds rate is well described as an I(1) process, forecasts of the rates themselves will tend to be very highly correlated; the correlation between the forecast *changes* in the Fed funds rate, $f_{1,t-1} - r_{t-1}$ and $\hat{r}_t^{VAR} - r_{t-1}$, will therefore be more informative. As reported in the first line of Table 2, the correlation between the forecast change in the Fed funds rate is 0.43, and the shock correlation is 0.35. Sampling uncertainty associated with the estimated VAR coefficients (readily apparent as "noise" in the forecast plotted in Figure 3) will increase the variance of the VAR forecasts, however, which will reduce the correlation between forecasts. One way to eliminate the effect of parameter uncertainty is to replace the variance of $\hat{r}_t^{VAR} - r_{t-1}$ in the denominator of the correlation coefficient with the variance of $f_{1,t-1} - r_{t-1}$. The result is just the $\hat{b}$ from the regression of the VAR forecast on the futures market's,

$$\hat{r}_t^{VAR} - r_{t-1} = a + b(f_{1,t-1} - r_{t-1}) + e_t .$$

Figure 3

Forecast one-month change in the Fed funds rate



Making this adjustment for parameter uncertainty yields a $\hat{b}$ of 0.57, further improving the CEE model's measured fit with respect to the futures market benchmark.

How do forecasts from a "no change" forecast, $\hat{r}_t^{NC} = r_{t-1}$ compare? The implied change from this forecast is, of course, zero. Consequently, the "no change" model's forecast of the change in the Fed funds rate is uncorrelated with everything, including the forecasts from the Fed funds futures rate and the funds rate itself (hence the zeros in the second line of Table 2). On this criterion, obviously, the VAR provides the better description of monetary policy. By contrast, the correlation between the "no change" forecast *errors*, $r_t - r_{t-1}$, and the futures market surprises, $r_t - f_{1,t-1}$, is 0.52 – much higher than the CEE model's. Judged on this criterion, therefore, the "no change" forecast describes monetary policy better than the VAR.

How can the "no change" forecast errors be more highly correlated with the Fed funds futures surprises than the VAR's, when the VAR's forecasts are closer to the futures market's? The answer, it turns out, is that the correlation between shocks says very little about how well the VAR describes monetary policy, and a lot more about small deviations from efficiency in the Fed funds futures market.

2.1 Anatomy of a correlation

The correlation between Fed funds futures surprises, u_t*, and forecast errors from an econometric model (e.g., a VAR), $\hat{u}_t$, can be written as:

$$\rho(\hat{u}_t, u_t^*) = \frac{\text{Cov}(u_t^*, \hat{u}_t)}{\sqrt{\text{Var}(u_t^*)\text{Var}(\hat{u}_t)}}$$

Since the variance of the Fed funds futures surprise is the same for each $\hat{u}$ we consider, differences in the correlation between shocks must be attributable either to differences in the covariance between the shocks, or to the variance of the estimated errors. Substituting $r_t - \hat{r}_t$ for $\hat{u}_t$, the covariance term in the numerator can be written as:

$$\text{Cov}(u_t^*, \hat{u}_t) = \text{Cov}(r_t, u_t^*) - \text{Cov}(\hat{r}_t, u_t^*)$$

or in terms of the change in r,

$$\operatorname{Cov}\left(u_t^*, \hat{u}_t\right) = \operatorname{Cov}\left(\Delta r_t, u_t^*\right) - \operatorname{Cov}\left(\hat{\Delta} r_t, u_t^*\right)$$

where $\hat{\Delta} r_t = \hat{r}_t - r_{t-1}$. Writing the covariance term in this way reveals two important features.

First, the covariance between the realized change in the Fed funds rate and the futures market surprise, Cov(Δr_t, u_t*), mechanically builds in a positive correlation between the two shocks. Just as important, this contribution to the covariance is wholly independent of the model's forecasts. In fact, it will be positive even if the model is of no use whatsoever in forecasting the Fed funds rate, as in the case of the "no change" model discussed above.

The second key observation is that a positive covariance between the model's forecast, $\hat{\Delta} r_t$, and the futures market surprise, u_t*, will *reduce* the covariance between the shocks. Market efficiency implies a zero covariance between u_t* and elements of the $t - 1$ information set. In practice, however, it is highly unlikely that the sample covariance will be zero even if the market *is* efficient. Indeed, Krueger and Kuttner (1996) found that this covariance, while nonzero, was generally statistically insignificant.

Taken together, these two observations explain the "no change" forecast's surprisingly high correlation with the futures market surprises. As reported in Table 3, the covariance between Δr_t and u_t* is 127.1, while the covariance between $\hat{\Delta} r_t$ and u_t* is identically zero. Dividing by the relevant standard deviations yields the correlation of 0.52 – well in excess of the CEE model's, despite the zero correlation between the forecasts themselves.

An analogous breakdown for the CEE model reported on the second line of Table 3 shows that a positive covariance between the VAR's predictions and the futures market surprises partially accounts for the model's small shock correlation. The relevant covariance is 39.5; subtracting this number from 127.1 and dividing by the relevant standard deviations yields the correlation of 0.32 (without a degrees-of-freedom adjustment). Had the Fed funds futures surprises been orthogonal to the VAR forecast, the correlation would have been 0.46.

Table 3

Components of the shock correlation

Model	$\rho(\hat{u}_t, u_t^*)$	$\operatorname{Cov}(\hat{\Delta} r_t, u_t^*)$	$\rho(\hat{\Delta} r_t, u_t^*)$	$\operatorname{Var}(\hat{u}_t)$
No change	0.52	0	0	18.7
CEE	0.32	39.5	0.14	21.1
T-bill	0.62	–34.9	–0.12	19.9
Modified CEE	0.37	24.7	0.09	21.3

Notes: The standard deviation of the Fed funds futures shock, $\sqrt{(\operatorname{Var}(u_t^*)}$, is 13.1, and the covariance between the change in the Fed funds rate and the Fed funds futures surprise, Cov(Δr_t, u_t^*), is 127.1. Units are basis points. Statistics are *not* adjusted for degrees of freedom.

One interpretation of this result is that the futures market is not efficient. The violation implied by this result is not quantitatively or statistically significant, however. The regression of the futures market surprise onto the VAR forecast has an R^2 of only 0.019, and the t-statistic on the CEE forecast's coefficient is only 1.41. But because standard deviations of the *shocks*, rather than the interest rate (or its change) appears in the denominator, a very small covariance can have a pronounced effect on the shock correlation.

The forecasts from a simple model involving the T-bill rate provide another illustration of the perverse properties of the shock correlation. A regression of the average Fed funds rate on two lags of the three-month T-bill rate,[9]

$$r_t = -0.41 + 0.79\ r_{t-1}^{tb} + 0.31\ r_{t-2}^{tb},$$

was used to generate (in-sample) one-month-ahead predictions. Since the T-bill rate presumably incorporates expectations of subsequent months' Fed funds rate, it comes as no surprise that this equation's forecasts are highly similar to those from the Fed funds futures market; in fact, the regression of the former on the latter gives a coefficient of 0.99. Yet the correlation between the shocks is 0.62 – only 20% higher than the "no change" forecast's. Again, the nonzero covariance between the model's forecasts and the futures market shock violates the assumption of strict orthogonality, but since this covariance is negative, it increases the shock correlation.[10] But the forecast's volatility is somewhat higher than the futures market's, and this reduces the correlation.

2.2 Does the VAR use too much information?

Aside from a violation of strict market efficiency, one reason for the CEE forecast's covariance with futures rate surprises is that the VAR incorporates "too much" information. The VAR forecast of the November funds rate (say) uses October's data, even though the most recent data on employment and prices is from September.[11] Moreover, much of these data are subsequently revised, and as Orphanides (1997) showed, the revisions can have a major impact on the fit of simple monetary policy rules. Consequently, the correlation between the futures market surprises and the VAR forecasts may be an artifact of the VAR's information advantage.

To see what impact this might have on the correlations, we re-ran the CEE equation with additional lags on payroll employment and consumer prices. (The reserves and money statistics are essentially known by the end of the month.) The results of this exercise appear in the final row of Table 3, labeled "modified CEE." This change reduces the positive covariance between the VAR forecasts and the futures market surprises somewhat, as would be expected if it were the result of the VAR's information advantage. Substituting unrevised, real-time data in place of the revised data used here might further reduce the covariance.

3. Parsimony and parameter instability

As shown above, the positive covariance between futures market surprises and VAR forecasts partially accounts for the low correlation between the VAR forecast errors and the futures market surprises. The dissection of the correlation coefficient also revealed a second culprit. the variance of the shocks from the VAR is considerably higher than the futures market surprises. Since the square root of this variance appears in the denominator, it, too, will reduce the measured correlation.

What accounts for the VAR's inflated shock variance? One possibility is the VAR's generous parameterization – 85 parameters in the monthly CEE specification. The profligacy of the CEE model surely explains the imprecision of the shock estimates, but it alone cannot explain the shocks' implausibly high variance, so long as the "true" model is nested within it.

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9 The regression uses last-day-of-the-month T-bill rate data, and it is estimated over the January 1961 through December 1997 sample.

10 The lagged futures rate is itself weakly (negatively) correlated with the futures market surprise.

11 For this reason, Krueger and Kuttner (1996) were careful to introduce additional lags when testing futures market efficiency.

Estimating the VAR over the 1961–97 sample might, however, contribute to the shocks' volatility if the parameters changed over time. A model estimated over a sample that included the 1979–82 M1 targeting regime, for example, will almost surely be inappropriate for later periods when the Fed's weight on monetary aggregates is smaller – if not zero. The spurious inclusion of M1 could therefore introduce noise into the forecasts for later periods. Other research has turned up significant time variation along these lines. Friedman and Kuttner (1996) estimated a time-varying-parameter version of a funds rate equation, and found significant variation in the coefficient on the monetary aggregates corresponding to the shifts in targeting regimes. Instability has also been documented by Pagan and Robertson (1995).

Table 4 reports the results of shortening the lag lengths and estimating the CEE model over shorter sample periods.[12] Comparing the twelve-lag to the six-lag results for the full sample shows that greater parsimony increases the shock correlation slightly, presumably by reducing the covariance between futures market surprises and the VAR forecasts. (Obviously, eliminating *all* right-hand-side variables drives this covariance to zero.) But greater parsimony actually *reduces* the correlation between the forecasts, and the forecast RMSE falls only slightly. The reduction in the number of coefficients to be estimated shrinks the standard error drastically, however.

Table 4

Improving the VAR forecasts

	Standard deviation of shocks	Forecast RMSE	Average standard error of shocks	Forecast $\hat{b}$	Shock correlation
One month					
Futures rate	13.1				
CEE model					
12 lags, full sample	23.5	27.1	20.6	0.57	0.35
6 lags, full sample	21.9	26.1	15.7	0.33	0.34
6 lags, post-83	16.2	25.6	6.5	0.52	0.54
Two months					
Futures rate	20.8				
CEE model					
12 lags, full sample	35.4	43.7	34.0	0.58	0.38
6 lags, full sample	34.1	42.0	24.2	0.39	0.39
6 lags, post-83	24.1	38.8	9.1	0.67	0.53
Three months					
Futures rate	29.6				
CEE model					
12 lags, full sample	48.2	60.7	42.1	0.55	0.47
6 lags, full sample	48.0	59.8	27.6	0.40	0.54
6 lags, post-83	31.5	51.5	10.7	0.75	0.60

Notes: The reported statistics are based on the May 1989 through December 1997 sample. The in-sample standard deviation and correlations are adjusted for the degrees of freedom used in estimation. Units are basis points.

The results improve considerably when the estimation period is restricted to the January 1983 through July 1997 sample. For one thing, the forecasts are now much less noisy. At 16.2 basis points, the standard deviation of the estimated one-month-ahead shocks is now only slightly larger than the futures market's.[13] The correlation between the shock measures rises to 0.54, but again the positive covariance between the model's forecast and the futures market surprises again prevents it

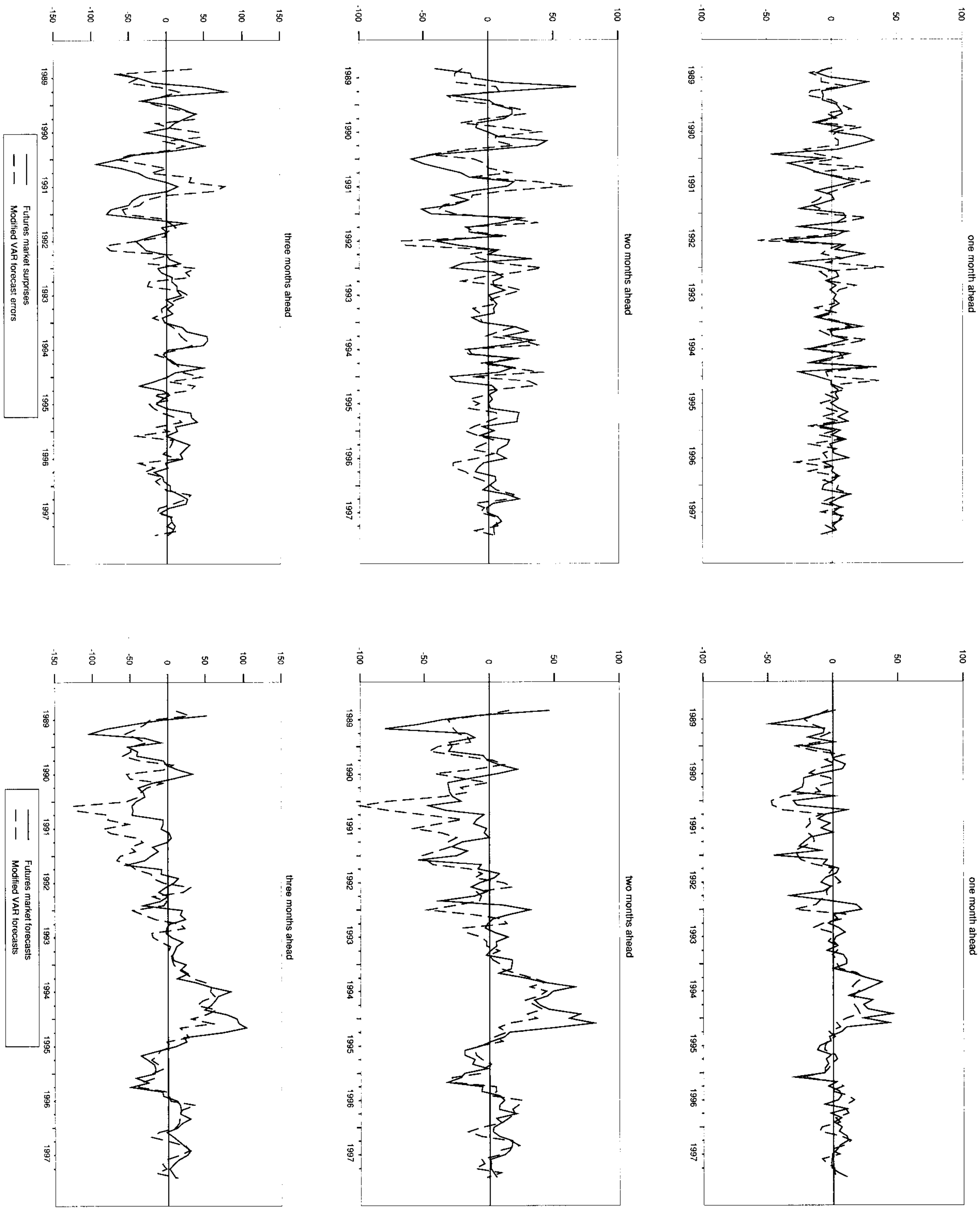
[12] As in Table 1, the correlations and standard deviations are adjusted for the degrees of freedom used in estimation.

[13] A similar result is apparent in Figure 7 of Christiano, Eichenbaum and Evans (1997).

from rising even higher. The regression of the model's forecast on the futures market's yields a coefficient of 0.52. The VAR does even better at longer horizons. At three months, the estimated $\hat{b}$ for the forecasts is 0.75, and forecast errors' correlation is 0.60. The forecasts and shocks plotted in Figure 4 confirm that the VAR approximately mimics the systematic and unsystematic changes in the funds rate implied by Fed funds futures rates.

Figure 4

Forecasts and errors from modified VAR equation



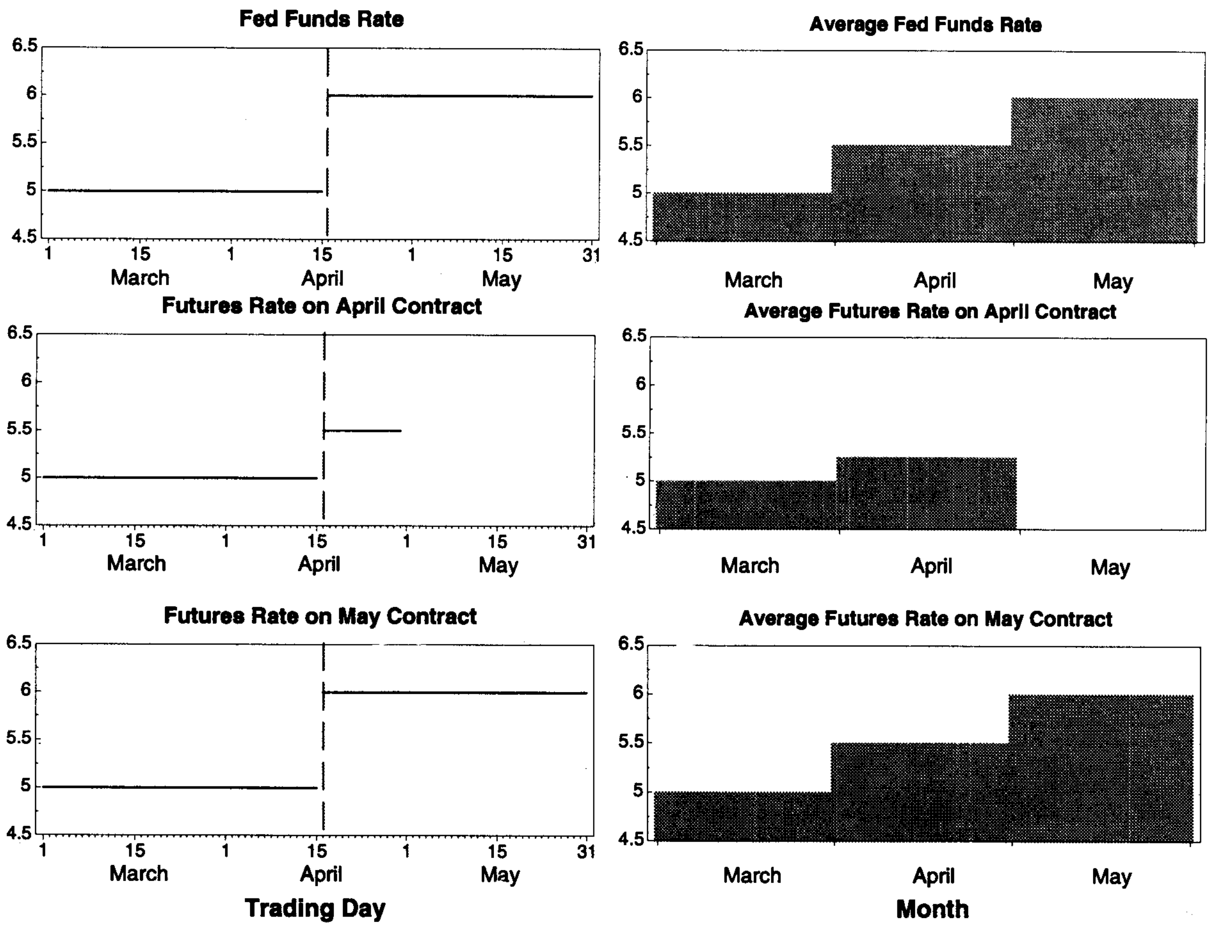
4. Time aggregation

The timing of the surprises extracted from monthly or quarterly VARs is, of course, somewhat ambiguous. At first glance, it would seem that shocks derived from the Fed funds futures rate would be free of such ambiguity. That turns out not to be true, however. The Fed funds futures contract's settlement price is based on the monthly *average* of the overnight Fed funds rate, which creates a time aggregation problem. Consequently, the timing and magnitude of policy shocks based on the futures rate are also ambiguous.

To illustrate this problem, consider the following scenario. The Fed funds rate is 5% in March, and this rate is expected to prevail through May. Now suppose that on April 16, the Fed unexpectedly raises the target Fed funds rate to 6%, and that the new rate is expected to remain in effect through May. Assume the Fed does, in fact, leave the rate at 6%. April's average Fed funds rate is 5.5%, reflecting 15 days at 5% and 15 days at 6%. The path of the Fed funds rate and the monthly averages are shown in the top row of Figure 5.

Figure 5

Financial markets' response to a Fed funds surprise



How will the futures rates respond to the surprise? With no change in the Fed funds rate expected, the futures rates corresponding to the April and May contracts will be 5% up through April 15. On April 16, the day of the surprise, the futures rate for the May contract will rise to 6%, reflecting the expectation that the 6% rate will prevail throughout May. But since the April contract is settled against April's *average* funds rate, the futures rate for the April contract will rise to only 5.5%. The paths of the futures rates are depicted in the left-hand column of the second and third rows of Figure 5.

The Fed's action on April 16 represents an unexpected increase of 100 basis points relative to expectations on April 15 and before. How should the Fed funds surprise be measured using monthly futures market data? Conceptually, this is simply the realized Fed funds rate minus its conditional expectation as measured by the futures data. But there are two complications. First, the futures contract is settled against the monthly average of the daily Fed funds rate; consequently, the realized Fed funds rate is taken to be the monthly average of the daily Fed funds rates. The second issue is whether to use point-in-time or average futures rate data in forming the conditional expectation.

Suppose we use the one-month-ahead futures rate on the last day of the previous month (e.g., the rate for the April contract as of March 31) as the conditional expectation, and measure the surprise relative to the monthly average Fed funds rate. In this example, summarized in the top panel of Table 5, the April surprise can be computed as April's average Fed funds rate (5.5%) minus the March 31 Fed funds futures rate for the April contract (5%), yielding only a 50 basis point surprise. Again using the futures rate from the last day of the previous month, the May surprise is calculated as the May average Fed funds rate (6.0%) minus the April 30 Fed funds futures rate for the May contract (6.0%), yielding no surprise. Recalling that the average funds rate increases 50 basis points in both April and May, the first 50 basis point increase is taken to be a surprise, while the second 50 basis points is anticipated. And yet, the example states clearly that the 100 basis point increase is a complete surprise on April 16. Last-day-of-month futures data, therefore, will tend to understate the magnitude of the true shock.

Table 5

Funds rate surprises using alternative measures of expectations

	Month		
	March	April	May
Average Fed funds rate	5.0%	5.5%	6.0%
Futures rate on last day of month	5.0%	6.0%	6.0%
implied Fed funds surprise	0	+50 b.p.	0
implied expected change	0	+50 b.p.	0
Average futures rate over month	5.0%	5.5%	6.0%
implied Fed funds surprise	0	+50 b.p.	+50 b.p.
implied expected change	0	0	0

An alternative way to measure the Fed funds surprise is to use the monthly average of a contract's futures rate, depicted in the right-hand column of the second and third rows of Figure 5, for the conditional expectation. This measure of the Fed fund surprise preserves the size of the shock's cumulative impact, but spreads it out over two consecutive months. As summarized in the bottom panel of Table 5, the April surprise is computed as the April average fed funds rate (5.5%) minus the March daily average of the April contract rates (5.0%), which is a 50 basis point surprise. The May surprise is the May average fed funds rate (6%) minus the April daily average of the May contract rates (5.5%), yielding a 50 basis point surprise. More generally, the surprise based on average futures rates will be a convex combination of the true shocks,

$$r_t - \bar{f}_{1,t-1} = \theta u_t + (1-\theta)u_{t-1}$$

which implies an MA(1) structure for the average shocks,

$$r_t - \bar{f}_{1,t-1} = (1+\phi L)e_t$$

where $\phi = (1 - \theta)/\theta$ and $e_t = \theta\, u_t$.[14] Econometric methods, like those of Hansen and Hodrick (1980)

[14] Estimating the MA(1) model gives $\hat{\phi} = 0.38$. This implies $\hat{\theta} = 0.72$, which is consistent with surprises typically occurring on the 8th day of the month.

and Hayashi and Sims (1983), exist to deal with the resulting moving-average error structure in market efficiency tests, but recovering the original "true" shock from the time-averaged data is generally not possible.

In the examples described so far, the monthly value of the Fed funds rate is taken to be the monthly average of the daily rates. How would the calculation be affected if the Fed funds rate from a single day were used in place of the monthly average? In this example, the April 30 Fed Funds rate (6.0%) minus the March 31 Fed funds futures rate for the April contract (5.0%) gives the correct 100 basis point surprise. However, other examples would generate problems with this calculation. Suppose that data were released in the first week of April that indicated the FOMC's normal response would be to increase the Fed funds rate by 50 basis points to 5.5%; and then on April 16 the actual policy move was 100 basis points to 6%. In this case, 50 basis points is anticipated, and the other 50 basis points is a surprise. But the calculation above is unaffected by the first week's data release, so the surprise is overstated by the amount of the mid-month's revision to anticipated policy. Finally, since the settlement price of the Fed funds futures contract is based upon the monthly average, there is little reason to believe the futures rates would satisfy conditions of unbiasedness and efficiency relative to the last-day-of-the-month Fed funds rate.

One way to reconstruct the "true" April shock is to rescale the first measure of the surprise. Specifically, compute the surprise as the April average Fed funds rate (5.5%) minus the March 31 Fed funds futures rate for the April contract (5%), and multiply the surprise by the factor m/τ, where m is the number of days in the month and τ is the number of days affected by the change. In the scenario described above, for instance, the measured surprise of 50 basis points is scaled up by a factor of two to yield the correct 100 basis point shock. This procedure only works when the dates of policy changes (and potential changes) are known, so it would only apply to the post-1994 period in which all changes in the target Fed funds rate occurred at FOMC meetings.[15] Prior to that time, most changes in the target occurred unpredictably *between* meetings. In this case, inferring the size of the "true" shock involves expectations of when the policy action occurs as well as the direction and magnitude of the change. This is beyond the scope of our analysis.

To get some sense of the quantitative importance of time aggregation, we computed the standard deviation of the rescaled Fed funds futures surprises for the post-1994 period, and compared it to the standard deviation of the unscaled shocks for the same period. The results appear in Table 6. As shown on the first line of the table, the volatility of the policy shocks rescaled in this way is dramatically higher – 28.2 basis points compared with 10.9 basis points for the unscaled shocks.

Table 6

Volatility of unscaled and rescaled Fed funds futures shocks

	Standard deviation	
	Unscaled	Rescaled
Using prior month's futures rate		
Average of effective Fed funds rate	10.9	28.2
Average of target Fed funds rate	8.9	19.7
Using spot month futures rate	–	22.9

Notes: The reported statistics are based on the February 1994 through December 1997 period. Units are basis points.

Rescaling the shocks in this way will exaggerate the effects of any transitory deviations of the funds rate from its target (i.e., "Desk errors"), however, so it will tend to overstate the volatility of the policy shocks. If there were an FOMC meeting two days before the end of the month, for example,

15 The only exception to this was the 25 basis point increase in the target in April 1994.

and if the monthly average Fed funds rate turned out to be 1 basis point above the target, the rescaling will result in a spurious 15 basis point shock. To reduce the effect of this noise, the average *target* Fed funds rate can be used in place of the effective rate. This procedure will distort the size of the shocks only to the extent that market participants expect the average effective rate to deviate from the target. Using the target rate, the standard deviation of the rescaled policy shocks is 19.7 basis points, compared with only 8.9 for the unscaled shocks.

An alternative way to gauge the effect of time aggregation on the magnitude of funds rate surprises is to use the "spot month" contract's price, which is based on the average Fed funds rate prevailing in the *current* month. Again, suppose that changes in the target Fed funds rate only occur immediately following an FOMC meeting, and, as in the example above, suppose that the FOMC meeting occurs on April 16. The difference between the April 16 and April 15 futures rates for the April contract would reflect the change in the expected path of the Fed funds rate over the April 16 to April 30 period. In the scenario described above, the spot month futures rate on April 15 would have been 5.0%, consistent with the "no change" expectation. On April 16, after the increase to 6.0%, the spot month futures rate would be 5.5%, since the contract's settlement price is based on an average that includes the first 15 days of the month, when the Fed funds rate was only 5.0%. As before, scaling the difference by the factor m/τ preserves the size of the shock. The result, as reported in Table 6, is a standard deviation of 22.6 basis points – very similar to the size of the shocks computed using end-of-month futures data and the target Fed funds rate.

Making adjustments for time aggregation results in futures-market policy surprises that are roughly twice as large as those that fail to make this adjustment. This result suggests that monetary policy is not as predictable as one might have suspected, and shows that time aggregation may distort comparisons between shocks based on futures rates and those from VARs.

Conclusions

Financial market data, such as Fed funds futures rates, are potentially useful benchmarks for evaluating econometric measures of systematic and surprise movements in monetary policy. The approach is not without its pitfalls, however. One hazard involves the interpretation of the correlation between Fed funds futures surprises and VAR shocks. This correlation contains little meaningful information relevant for assessing the VAR's description of monetary policy. As shown above, small deviations from the orthogonality condition implied by market efficiency can have a big effect on the correlation.

This is not to say that VAR's description of monetary policy is perfect. Their forecasts are imprecise and noisy, and there is some evidence to suggest parameter instability. Shorter lag lengths and a more judicious selection of starting date can mitigate these problems, however, and the results presented here suggest more research along those lines is warranted.

One important complication arising in comparisons between VARs and futures-market forecasts is time aggregation. This problem can distort the timing and magnitude of the estimated policy surprises: point-in-time futures rate data gets the timing right, but attenuates the magnitude, while average data gets the magnitude right but distorts the timing. This observation has important implications for attempts to draw inferences about the size of policy shocks from futures market data.

While the distortion created by time aggregation may have significant effects on the contemporaneous correlation between shocks, it is unlikely that it would affect the economy's estimated response to those shocks. Because an impulse response function can be thought of in terms of a regression of the relevant variable on a set of mutually uncorrelated shocks, merely shifting the shocks' dating a month – or even a quarter – in one direction or another may alter the *timing* of the response but have little effect on its shape or size. Consequently, the timing ambiguities identified above are probably irrelevant for measuring the real effects of monetary policy.

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Comments on: "Can VARs describe monetary policy?" by Charles L. Evans and Kenneth N. Kuttner

by David Tessier

This paper addresses an important question for monetary policy analysis, that is the usefulness of VARs to isolate monetary policy shocks. The authors overcome the identification problems inherent to structural VARs by focusing explicitly on the reduced form. Hence, they estimate a VAR and then use the error term from the Fed fund rate equation as a measure of the monetary shock. In order to assess the appropriateness of these shocks, they compare them with the surprises from the Fed fund futures market that are seen as the "true shocks". The estimated correlation between the shocks is quite low and this is not surprising given the VAR's poor performance, both in and out of sample. But the authors go a step further and give convincing arguments against using a correlation metric to evaluate the validity of shocks from VAR models. Then they conclude that VARs, although subject to some well-known pitfalls, are still valid for policy analysis.

The core of the argumentation assumes that the surprises from the Fed fund futures market are a valid benchmark. This could be questioned given the recent paper by Robertson and Thornton (1997), who note some identification problems in estimating expectations from the Fed fund futures market. Notwithstanding this caveat, I would argue that the conclusion of the authors might be reinforced by improving the specification of the VARs and consequently, the accuracy of predictions necessary to recover the shocks. There are two main sources of potential improvement. First, the authors specify a VAR model in which each equation contains 84 parameters, which is far from parsimonious. A more parsimonious representation could be obtained by applying a Bayesian specification or a "top-down strategy" to remove the non-significant parameters (see the Section 5.2.8 in Lütkepohl (1993)).

Second, the authors estimate their model with variables in levels and we know that the usual asymptotic results do not hold in that case, owing to non-stationarity and/or the absence of cointegration. As shown convincingly by Phillips (1998), ignoring these two characteristics may lead to less accurate predictions. In this paper, Phillips develops an asymptotic theory for forecasting and policy analysis that allows for nonstationary elements. Based on simulation results, he concludes that the data-determined reduced rank regressions and the error correction models produce better forecasts than unconstrained VARs (see the Section 4.3).[1] Closely related to this is a paper by Christoffersen and Diebold (1997) studying the impact of cointegration on short-term dynamics. They conclude that "ironically enough, although cointegration implies restrictions on low-frequency dynamics, imposing cointegration is helpful for short- but not long-horizon forecasting, in contrast to the impression created in the literature".

In that context, VAR modelisation, with appropriate specification procedures taking into account the problems involved in nonstationary systems, might indeed remain a useful and tractable tool for assessing the impact of monetary policy.

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[1] These results hold under the hypothesis that there really exists a problem of reduced rank regression. But given the number and the choice of variables retained by the authors, the presence of cointegration is highly plausible.

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Incorporating credibility in forward-looking models: some examples with QPM

Dinah Maclean*

Introduction

Countries are increasingly implementing monetary policy in a manner designed to build credibility. While the empirical evidence on whether or not credibility is greater is still inconclusive, changes in credibility could potentially have far-reaching consequences for the dynamics of the economy. The most obvious changes directly involve inflation expectations. If credibility is high, a reduction in the inflation target will quickly be reflected in expectations. This in turn will help bring inflation down to the new target, reducing the output losses associated with a disinflation. Similarly, if expectations are anchored by policy, the monetary authority will not have to react as aggressively to temporary inflation shocks. In addition, with increased credibility, agents will likely be willing to place greater weight on longer term forecasts. This in turn may be associated with increased contract lengths. All these changes may alter the speed with which policy can influence inflation, and thus the manner in which policy should be conducted.

Given the importance of issues surrounding credibility, explicit credibility effects need to be incorporated into models used for policy analysis and projections. This paper describes the development and calibration of a credibility effect using the Bank of Canada's Quarterly Projection Model (QPM). A methodology for incorporating credibility effects is developed, which uses a "perceived target" which captures agents medium-term inflation expectations. The model is calibrated taking a cautious approach, given the lack of conclusive empirical evidence which can be used to benchmark the degree of credibility. The perceived target is calibrated to ignore short-term changes in inflation, but to look forward to the medium term inflation profile, but it is given a relatively low weight in the overall calibration of expectations. Given the absence of other benchmarks, some of the key stylized facts to which the model is calibrated are left unchanged, such as the cumulative output gap associated with a disinflation. Nevertheless, the new model does incorporate changed characteristics. In particular, it decreases the variability of inflation in non-policy shocks, and reduces the degree of response needed by the monetary authority to offset such shocks.

The calibration presented in this paper should be viewed as an initial attempt to incorporate credibility effects. Above all, it highlights the need for further empirical work on such areas as whether expectations formation has undergone a structural change in the 1990s associated with credibility, and whether or not key stylized facts such as sacrifice and benefit ratios have changed. It will also be important to simulate the new model using real data. An interesting check of the current calibration will be to simulate the model over history and see how changing the weight on the perceived target alters the extent to which model expectations track other measures of expectations such as those coming from the Conference Board Survey of Forecasters.

The next section provides a brief overview of the nature of credibility and the expected changes associated with increased credibility. Section 2 gives an overview of QPM, particularly those aspects of the model which affect the degree to which the inflation target is incorporated into expectations, and thus the implicit level of credibility. Section 3 considers two alternative ways of more explicitly incorporating credibility effects into price expectations. Section 4 both reviews empirical evidence and presents two calibrations of QPM.

* The views expressed in this paper are solely those of the author. No responsibility for them should be attributed to the Bank of Canada.

1. Credibility

Credibility is generally interpreted as the extent to which agents believe the policy maker. In Canada, where the Bank of Canada and the Minister of Finance began announcing inflation targets in 1991, credibility can be defined as the extent to which people believe inflation will remain within the target bands.

We distinguish between two aspects of credibility: the "announcement" effect and the process of learning or gaining credibility over time. The announcement effect captures the extent to which a policy announcement immediately changes people's expectations and behaviour. This announcement effect will increase as the monetary authority gains credibility. If credibility in the targets has been growing in Canada in the 1990s, for example, then an announced change in the target should alter agents' inflation expectations. The size of the announcement effect is determined by the degree to which credibility is transferable from one policy to another. Moreover, an announcement effect may occur whether or not the announcement of a policy change coincides with a consistent policy action by the central bank. The second aspect of credibility deals with how quickly credibility is gained or lost over time, or the learning process during which agents alter the extent to which they incorporate the policy into their expectations. Over time the monetary authority may build up or lose credibility, depending on its success in achieving the targets.

Both the announcement effect and the process by which agents incorporate a target into expectations over time will be dependent on the dynamics of the economy. For example, in an economy where it is known that the monetary authority has little impact on inflation in the near term, it would be unreasonable for near-term expectations to adjust immediately to a new target. Similarly, if it is known that adjustment is costly and prices are sticky, this will affect the speed with which people expect inflation to move to the target, even when credibility in the monetary authority's commitment to the policy is high.

Changes in credibility can potentially have a large impact on the dynamics of the economy, and the extent to which various shocks are reflected in inflation. In general these changes are likely to make it easier for a monetary authority to maintain a stable rate of inflation, or implement a change in policy regime. Consider, for example, the case of a policy change such as a decrease in the inflation target. If credibility in the current policy is high, this will likely carry over to credibility in the new policy. Expectations will incorporate the new target much more rapidly than if expectations are more backward-looking. With faster adjustment of expectations, actual inflation will fall more quickly and/or the monetary authority will need to act less in order to achieve the new target. This in turn implies that the cumulative output loss associated with a disinflation, will be less with higher credibility.

Increased credibility has similar implications for a non-policy shock. In the case of an inflationary shock, for example, such as a depreciation in the exchange rate or a positive shock to demand, credibility will help to anchor inflation expectations. Expectations will not vary as much when people anticipate that the monetary authority will act to bring inflation back to the target. As a result, the monetary reaction does not need to be as great. The cumulative output losses in a deflationary shock will tend to be greater than in the case of less credibility, since less of the shock is offset directly by policy. Similarly, cumulative output gains from a positive shock will be greater.

While there is considerable agreement about the gains from greater credibility, far less is known about the process of building credibility. A wide variety of factors that likely influence credibility have been identified in the literature, such as the clarity of the goals of monetary policy, the extent to which policy is understood, the central bank's degree of autonomy and accountability, and the consistency of monetary and fiscal policies. There is no precise mapping, however, between different settings of say fiscal policy or institutional arrangements and a specific degree of credibility. Nevertheless, given the potential benefits from increased credibility for the monetary authority, an increasing number of countries are implementing monetary policies designed to increase credibility. Since increased credibility could have potentially large implications for adjustment in the economy, it is also important to incorporate credibility into economic models. The next section provides a brief overview of the Bank of Canada's Quarterly Projection Model (QPM) and in particular the expectations within the model. The remaining sections then look at ways of incorporating specific credibility effects, and review issues associated with the calibration of credibility effects.

2. Credibility and QPM

QPM is a forward-looking model, with expectations which are partly backward-looking, and partly based on model consistent values. While there are no explicit credibility effects within QPM, credibility is still embodied in the model as the degree to which the inflation target is imbedded in expectations. Initially, this comes through the model consistent component of expectations, and subsequently becomes embodied in the backward-looking component. This behaviour depends on such things as the exact structure of expectations formation and the way in which expectations feed through into the price equations. The degree of credibility is also embodied in many properties chosen in the overall calibration of the model and thus affects the model's response to a range of shocks. To assess the appropriateness of any approach, therefore, requires that a broad range of model properties needs to be considered. This section provides a short overview of QPM then considers the main parts of QPM which affect or incorporate assumptions about credibility. It also provides an illustration of a model change which intuitively may seem consistent with increasing the degree of credibility in the model, but does not produce model properties which are consistent with theory.

QPM consists of two models: a well-defined, neo-classical steady-state model and a dynamic model which traces the adjustment path between the starting conditions and the steady-state. There are three key groups of agents who determine the steady state conditions: consumers, firms and government. Consumers have a desired level of wealth and make decisions on savings and consumption over time to reach that level. Their behaviour is modelled on the Blanchard-Weil model of overlapping generations. Firms determine the capital stock and associated rates of investment. The government sector determines the level of debt and associated levels of government expenditure and taxation. These stocks, together with the rest-of-world economy, determine the level of net foreign assets. The exchange rate adjusts so as to ensure the current account balance is consistent with the flows needed to service any foreign debt. QPM is based on the "Almost Small Economy Assumption", which means that given unchanged conditions in the rest of the world, in order to export a greater quantity, the price of exports must fall.

There are three main sources of dynamics which determine the short and medium term adjustment in QPM: intrinsic dynamics, expectational dynamics and policy adjustment. The intrinsic dynamics capture the idea that adjustment is costly and therefore occurs over time rather than all at once. In making decisions, agents must balance the cost of being away from their desired levels against the costs of adjusting variables. Agents follow decision rules where variables such as prices are combinations of both backward and forward-looking elements. The forward-looking elements are based on expectations equations which are also a combination of backward looking and model consistent values, and in some cases steady-state variables.

There are two policy equations within QPM. A fiscal policy rule determines government expenditure and taxation based on an exogenously determined path for the debt-to-GDP ratio. There is also a monetary policy rule, where the monetary authority reacts to shocks by altering nominal short-term interest rates in order to bring inflation back to the target in the medium-term. The monetary policy rule is written in terms of the yield curve gap, or the difference between the slope of the nominal term structure and its risk-adjusted steady state value. The desired value of the yield curve gap is a function of a lagged value and the deviation of year-on-year inflation in consumer prices excluding food and energy (*CPI*) and the target rate of inflation, 6 to 7 quarters in the future:

$YCG = al*(YCG)_{t-1} + ((CPI_{t+6} - CPITARG_{t+6})+(CPI_{t+7} - CPITARG_{t+7}))/2$

where YCG is the yield curve gap and $(CPI_{t+i} - CPITARG_{t+i})$ is the difference between the year-on-year increase in the CPI and the annual inflation target i periods ahead.

2.1 Calibrated properties – Sacrifice and benefit ratios

The dynamic structure of QPM is not directly estimated; rather it is calibrated to reflect empirical evidence and established stylized facts. For example, currently, the model is calibrated to ensure that there is sacrifice ratio of 3 in a disinflation shock (i.e. in a 1% disinflation shock, the cumulative output gap is 3%), and a benefit ratio of 1 in an inflation shock. In contrast to the disinflation and inflation shocks where the sacrifice ratios are always restored, changes to the model often result in changes to cumulative output gains and losses in other shocks. The cumulative output losses and gains from other shocks are not specifically calibrated and there is generally less empirical evidence to help determine what they should be.

The properties of the disinflation and inflation shocks are based on non-linear Phillips Curve estimations by Laxton, Rose and Tetlow (1993) for the period 1975 to 1991. As mentioned above, the degree of lost output that the monetary authority needs to generate to offset temporary shocks to inflation or to implement policy changes, is greatly affected by the degree of credibility of policy. The current calibration, therefore, reflects the "average" degree of credibility in the period over which the equation was estimated. It is hard to assess how high this average credibility is. The rising rates of inflation over the 1970s likely eroded credibility while the relatively more stable and lower rates in the 1980s likely restored some of the lost credibility. In the 1990s, the announcement of inflation targets and the success in keeping inflation within the range is likely to have raised credibility further.

Any attempt to incorporate increased credibility, must include a re-assessment of the sacrifice and benefit ratios. In particular, if credibility is incorporated in such a way that the current target helps to anchor expectations, but changes in the target are not immediately believed, the cumulative output losses/gains in shocks where the target remains unchanged will alter, but the sacrifice ratio in a disinflation shock need not fall. Alternatively, if credibility effects also include rapid assimilation of a newly announced target into expectations, the sacrifice and benefit ratios in the disinflation and inflation shocks may fall. Choosing new values for these is problematic, however, particularly given the lack of any empirical evidence supporting a decline in the sacrifice ratio in Canada.

2.2 Expectations formation

Price expectations in QPM are in terms of price levels rather than inflation. Price levels are a combination of a backward and a forward-looking component. For example, the expected level of the log of consumer prices excluding food and energy in period T ($LCPI\_ET$) is the combination of a backward element including 5 lags of the past price level, and the model consistent value for period T:

$LCPI\_ET = BW*Backward + MC*Modelconsistent(T)$

where BW is the weight on the backward-looking component and MC is the weight on the model consistent component.

Currently, the weights on the backward and forward components of consumer price expectations are 0.8 and 0.2 respectively. The weights chosen for the backward and model consistent components obviously have a major impact on the speed with which expectations change given a change in the target, and the extent to which shocks to inflation flow through into expectations. Since the monetary policy rule is acting to move inflation in the medium term to the target, model consistent expectations in future periods increasingly incorporate the target rate of inflation.

2.3 Incorporation of expectations into prices

The effects of varying the weights on the backward and forward components of expectations are partly a function of the way in which expectations are incorporated into the price equation. Expectations feed into prices through the forward-looking component of the price equation, which is a weighted average of expectations of both marginal cost and the price up to eight periods ahead. For example, the forward-looking component of the *CPI* is:

LPCPIFOR = F(LCPI_E0...E8, LMC_E0...8)

where *LPCPI_ET* is the expectation of *LPCPI* in period *T* and *LMC_ET* is the expectation of the log of marginal cost in period *T*.

The expectations are weighted so that the most weight is placed on the near-term expectations, and the weights decline the further into the future are the expectations. This reflects the assumption that people are less certain about expectations of events further into the future, and therefore, base their behaviour more on shorter term expectations.

2.4 Examples of model properties – a disinflation and a demand shock

Figures 1 and 2 show the adjustment of key variables in a 1% disinflation (i.e. a reduction in the inflation target of 1 percentage point) and a 1% negative demand shock (i.e. total demand is reduced by 1 percentage point). In a disinflation, the model is calibrated to ensure that the monetary authority must raise real short-term interest rates by around 100 basis points in the first year. In nominal rates, this corresponds to an increase of 80 basis points in nominal short-term interest rates. The tighter yield curve acts directly to reduce consumption and investment. Higher interest rates also lead to an appreciation of the exchange rate, which further reduces inflationary pressures. CPI inflation falls 0.13 percentage points by the end of the first year, and takes 4-5 years to reach the new target. As mentioned above, the coefficients of the output gap within the price equations are adjusted to ensure a cumulative output gap of 3%.

In a negative demand shock, the monetary authority responds by reducing short-term interest rates in order to loosen the slope of the yield curve. Nominal short-term interest rates fall 80 basis points in the first year. CPI inflation troughs at a rate 0.4 percentage points below control. Output returns to control at the end of the second year following the shock, but there is some secondary cycling for the next few years. The cumulative output gap is just over 1%.

2.5 Increasing the model consistent component of expectations

Increasing the weight on the forward-looking or model consistent component at first seems very close to increasing the degree of credibility in the model, and gives results consistent with increased credibility when considering a change in the inflation target. In the case of a non-policy shock such as a demand shock, however, increasing the weight on the model consistent component gives results which are the opposite of those expected from increased credibility.

The effects in a disinflation shock of increasing the weight on the model consistent component of expectations are illustrated in Figure 3. Figure 3 shows the effects of a 1 percentage point decline in the inflation target, with a range of models where the weight on the model consistent component of expectations for both the CPI and the GDP deflator range from 0.1 up to 0.7. (Note that the legend only shows every other line.) The models with a larger weight on the model consistent component show a more rapid decline in inflation in a disinflation shock, since the forward-looking component pulls expectations down while the backward component slows the adjustment in expectations. Similarly, the monetary authority does not have to tighten the yield curve gap (the policy instrument) as much as when greater weight is placed on backward-looking expectations and the resulting cumulative output loss is reduced. In this case, the results from equations where more weight is placed on the model consistent component are equivalent to those expected from a greater

Figure 1

1% disinflation

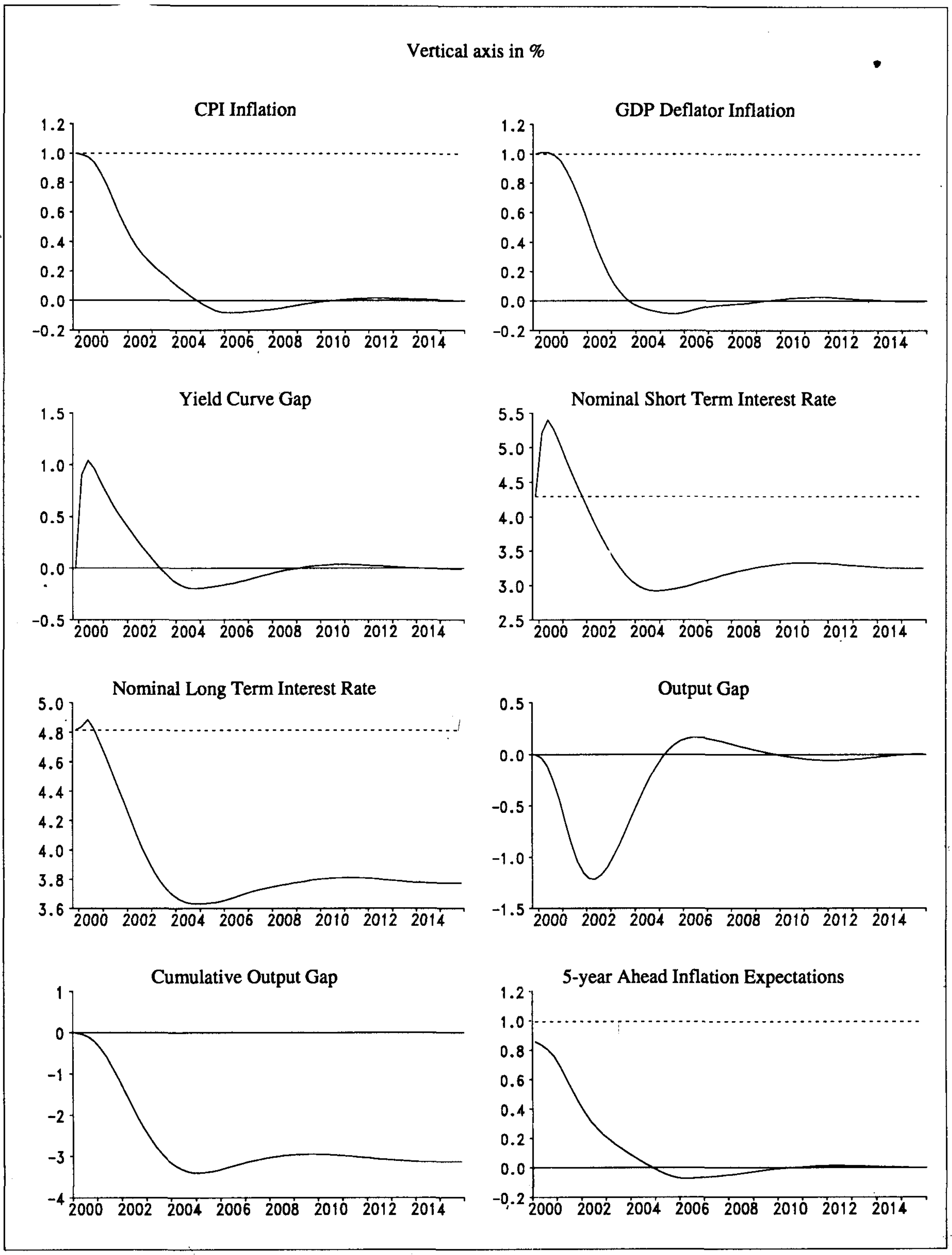


Figure 2

1% negative demand shock

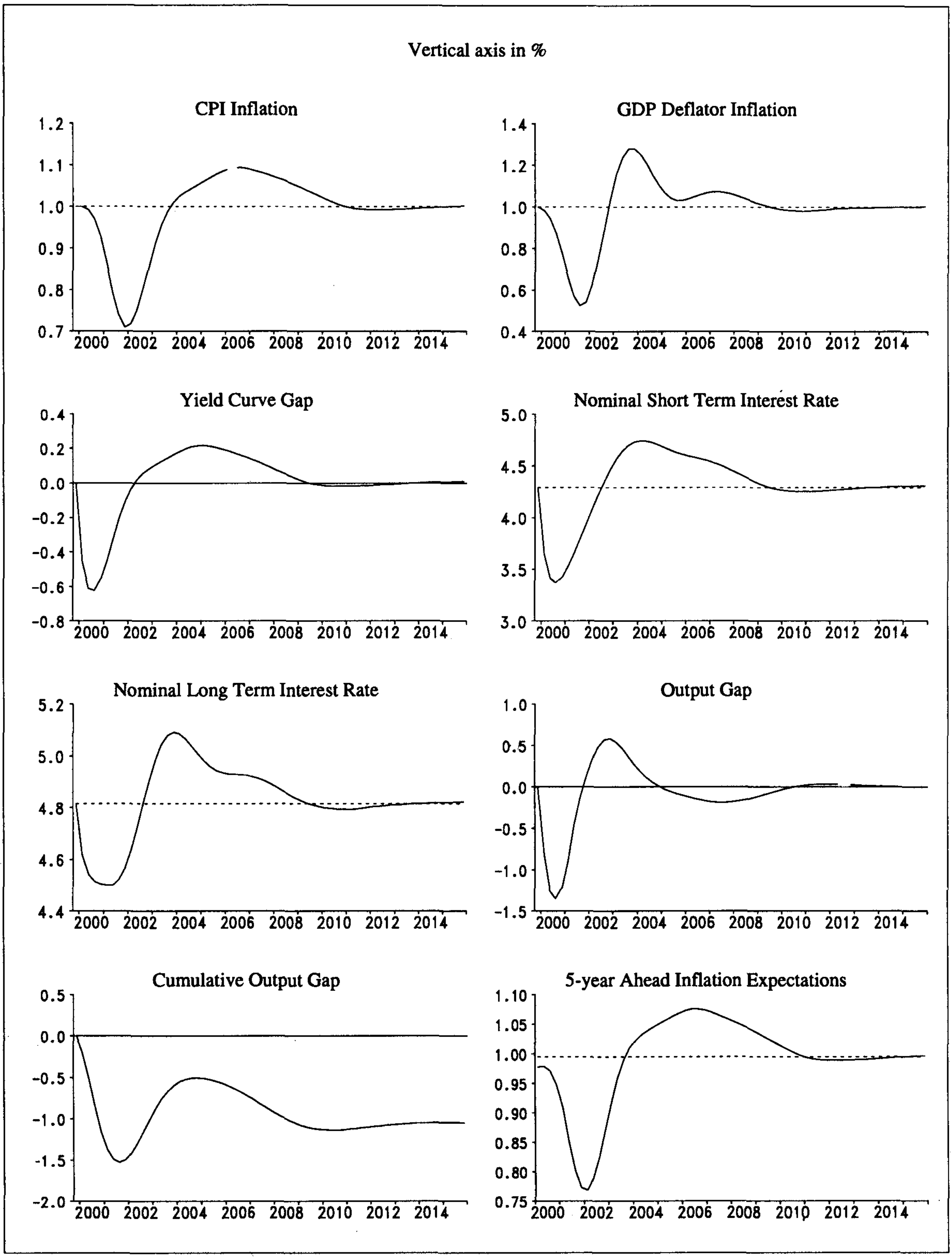


Figure 3

1% disinflation – varying weights on model consistent component

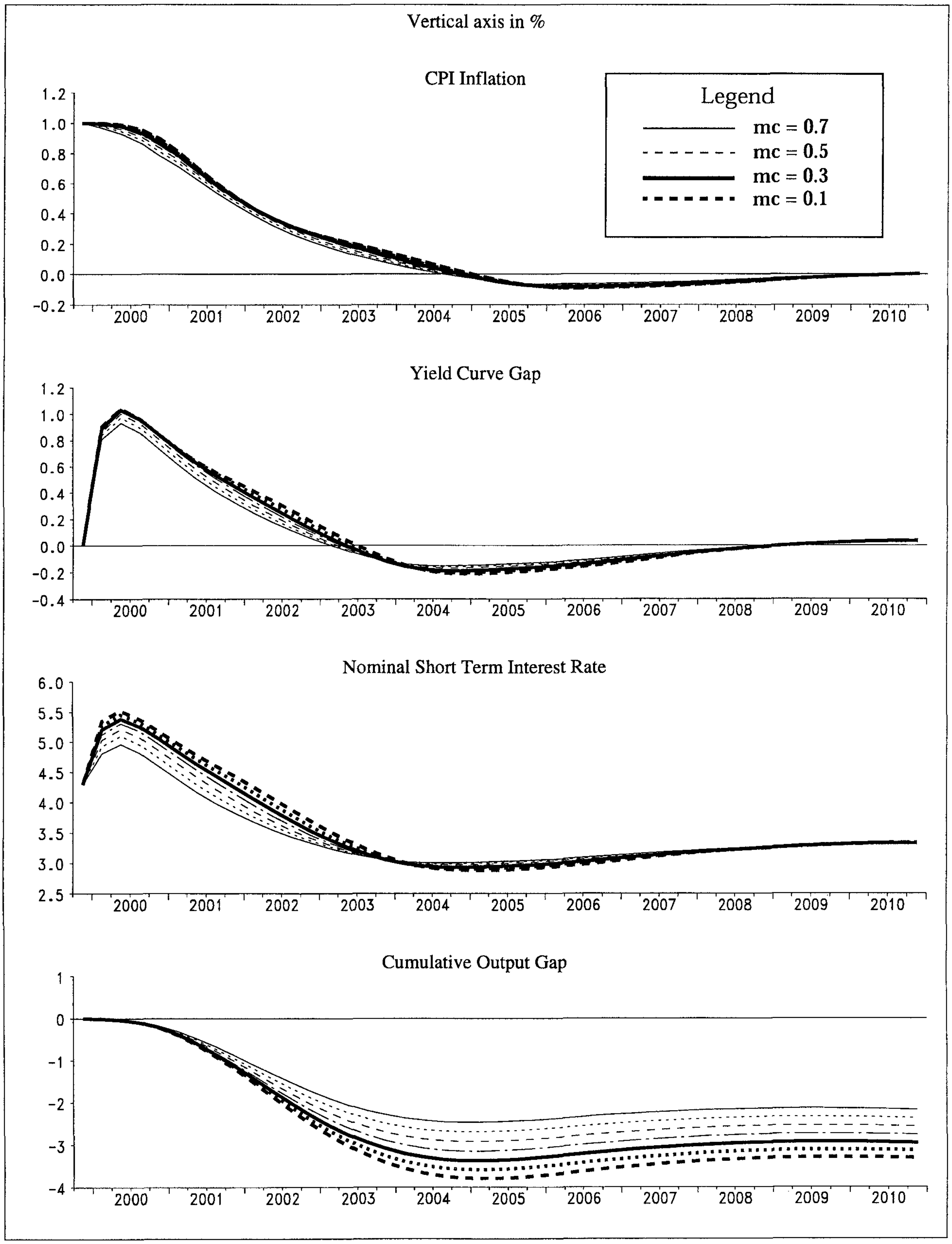
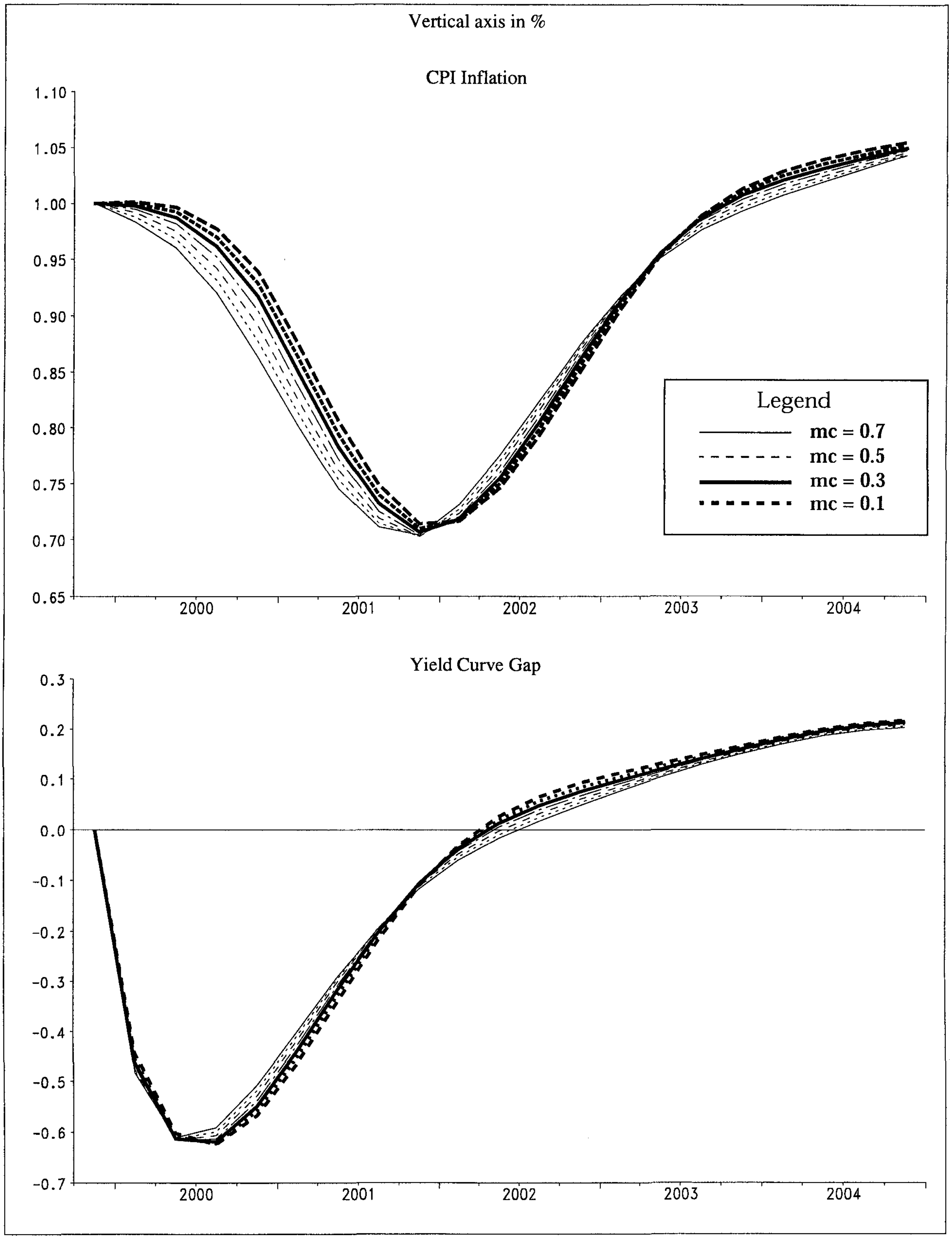


Figure 4

1% negative demand shock – varying weights on backward component



degree of credibility – faster adjustment of expectations and a smaller output loss associated with moving to a lower target.

Figure 4 shows for the same models, the effect of a 1% decline in demand. In the standard shock, a decrease in demand leads to excess supply and puts downward pressure on inflation. Expectations of inflation also fall putting further downward pressure on the price level. Initially, the fall in expectations comes through the model consistent component, then over time as inflation falls, through the backward-looking component. The changes in expectations feed directly through into inflation. The monetary authority responds by decreasing interest rates which boosts consumption and investment. Decreased interest rates also lead to a depreciation in the exchange rate which helps generate demand for exports and import-substitution, as well as having direct pass-through effects on the price level.

In theory, where there is increased credibility, inflation expectations should act as more of an anchor in such a shock, since people believe the monetary authority will act to bring inflation back to the target. Having less of a fall in expectations will put less downward pressure on inflation and will reduce the degree to which the monetary authority must act. Figure 4 shows, however, that the opposite is true when the weight on the model consistent component is increased. In the first few years of a demand shock, the model consistent component actually pulls expectations away from the target whereas it is the backward component which provides inertia. This is because the value generated by QPM for inflation in a specific quarter, puts greater weight on short run model dynamics, since greater weight is placed on near-term expectations. Agents are responding more to the short run dynamics of the shock, thereby causing greater, not less variation in expectations. Increasing the weight on the model consistent component, therefore, leads to a greater trough in inflation and more need initially for the monetary authority to offset the shock. The greater the weight on the model consistent component, the smaller the cumulative output gap in a negative demand shock, since the monetary authority offsets more of the negative shock to output.

It is also important to note that changing the calibration of expectations, changes the monetary response needed for a given shock, and thus the output losses and gains associated with that shock. Moreover, as described above, theory suggests that greater credibility should be associated with less need for the monetary authority to respond to a shock and decreased costs of disinflating. This implies that incorporating a greater credibility effect in a model, should be accompanied by a decrease in the sacrifice ratio. Unfortunately, however, given that the sacrifice ratio is a result of the calibration of the model, it is not obvious how to determine a new benchmark for the value of the sacrifice ratio, since empirical work has not yet shown any evidence of declining inflation/output tradeoffs.

3. Incorporating the target into expectations

The goal of monetary policy in QPM is to keep inflation in the CPI close to the target. The easiest way of incorporating credibility effects more explicitly in QPM, therefore, is to introduce the target into price expectations. As shown below, though, this also introduces an announcement effect which is both inconsistent with the intrinsic dynamics within QPM and an undesirable feature for a projection model. An alternative formulation is tried, therefore, where expectations include a perceived target. This reduces the problems associated with the announcement effect and also provides greater flexibility for calibrating the model. The perceived target is based on the formulation of expectations in Black and Rose (1997).[1]

[1] See also Black, Macklem and Rose (1998).

3.1 Introducing the target

Price expectations are a combination of a backward component and the model consistent value. The target was initially incorporated by introducing a third term – the price level implied by the inflation target.

The equation for expectations of the log level of the CPI (*LCPI*) in period *T* is:

$LCPI\_ET = XBW*DC*(BACKWARD) + XMC*LCPI(T) + (1\text{-}XMC\text{-}XBW*DC)*LCPITAR\_ET$

where *XBW* is the weight on the backward component, *XMC* the weight on the model consistent value, (1-*XMC*-*XBW***DC*) is the weight on the target, and *LCPITAR_ET* is the price level implied by the target in period *T*. *DC* is a decay factor which takes a value between zero and one. The value of *DC* can be fixed, implying a fixed weight on the target, or it can decline over time implying a greater weight on the target for longer term expectations.[2] This discussion focuses on the results of the constant decay model.

LCPITAR_Ei becomes a new endogenous variable, which is created using the specified year-over-year inflation target. The expected target level in the first period is the previous period's price level, plus the quarterly rate of change implied by the inflation target:

$LCPITAR\_E1 = LCPI(\text{-}1)+LOG(CPITAR(i))$

Expectations for the target level in subsequent periods are calculated as the target level expectation of the previous period plus the quarterly rate of change implied by the inflation target:

$LCPITAR\_Ei = LCPITAR\_E(i\text{-}1)+\mathrm{LOG}(CPITAR(i))$

Expectations of the target were incorporated into both expectations of the GDP deflator and expectations of the CPI.

To explore the implications of introducing the target directly in price expectations, two shocks were introduced: a disinflation shock, and a negative demand shock. The models were simulated with three different values of the decay parameter: 0.95, 0.8 and 0.6. The results are shown in Figures 5 and 6.

The results for the horizon-dependent model are shown in Appendix 1.

In the disinflation shock, the greater the weight on the target, the larger the initial announcement effect. This can be seen in the five-year ahead inflation expectations which start out at a much lower rate than the base model, the higher the weight on the target. The immediate impact that a change in the target has on expectations is also very evident in the initial decline in nominal long-term interest rates. The decline in expectations in turn puts downward pressure on the actual rate of inflation. In the first quarter of the shock, when the decay is set at 0.95 (i.e. only a small weight on the target) CPI inflation falls 0.02 percentage points, compared to 0.08 percentage points with a decay of 0.6. In the base model there is essentially no announcement effect in the first quarter. By the fourth quarter, CPI inflation has fallen 0.27 percentage points with a 0.95 decay, compared to 0.4 percentage points (almost half of the movement to the new target) in the model with a decay of 0.6, and 0.13 percentage points in the base model.

Earlier, credibility effects were divided into announcement effects, and the speed which which agents "learn" over time. In QPM, however, these learning effects are not very clear, because the path expectations take over time is greatly affected by the reaction of the monetary authority. This is evident by comparing the time it takes the five-year ahead expectations of inflation and actual inflation to reach the target. Expected inflation over the first four years is lower the greater the weight

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[2] This is the form currently used for expectations which decay to the steady state, for example, expecations of output and the real exchange rate.

Figure 5

1% disinflation – fixed coefficients

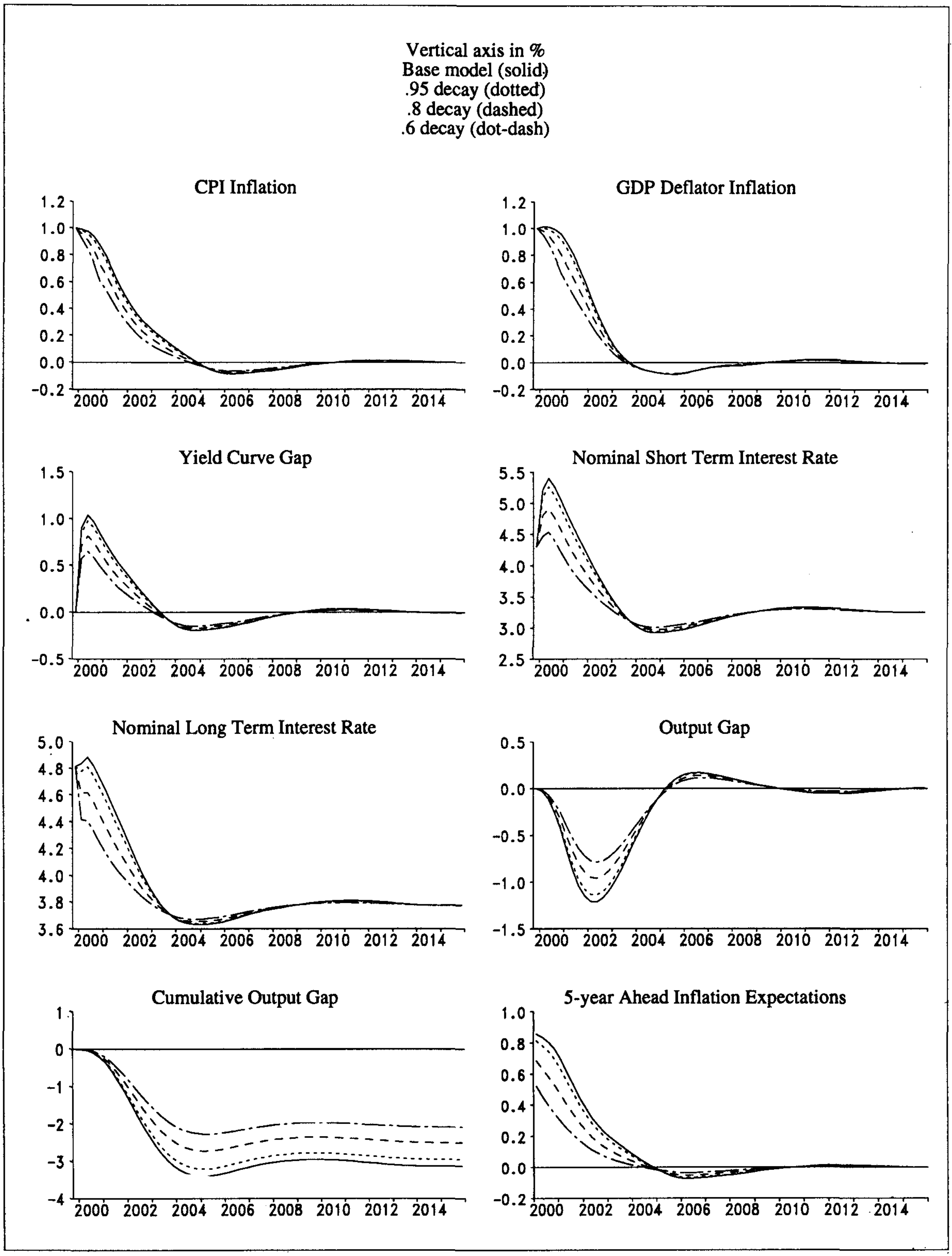
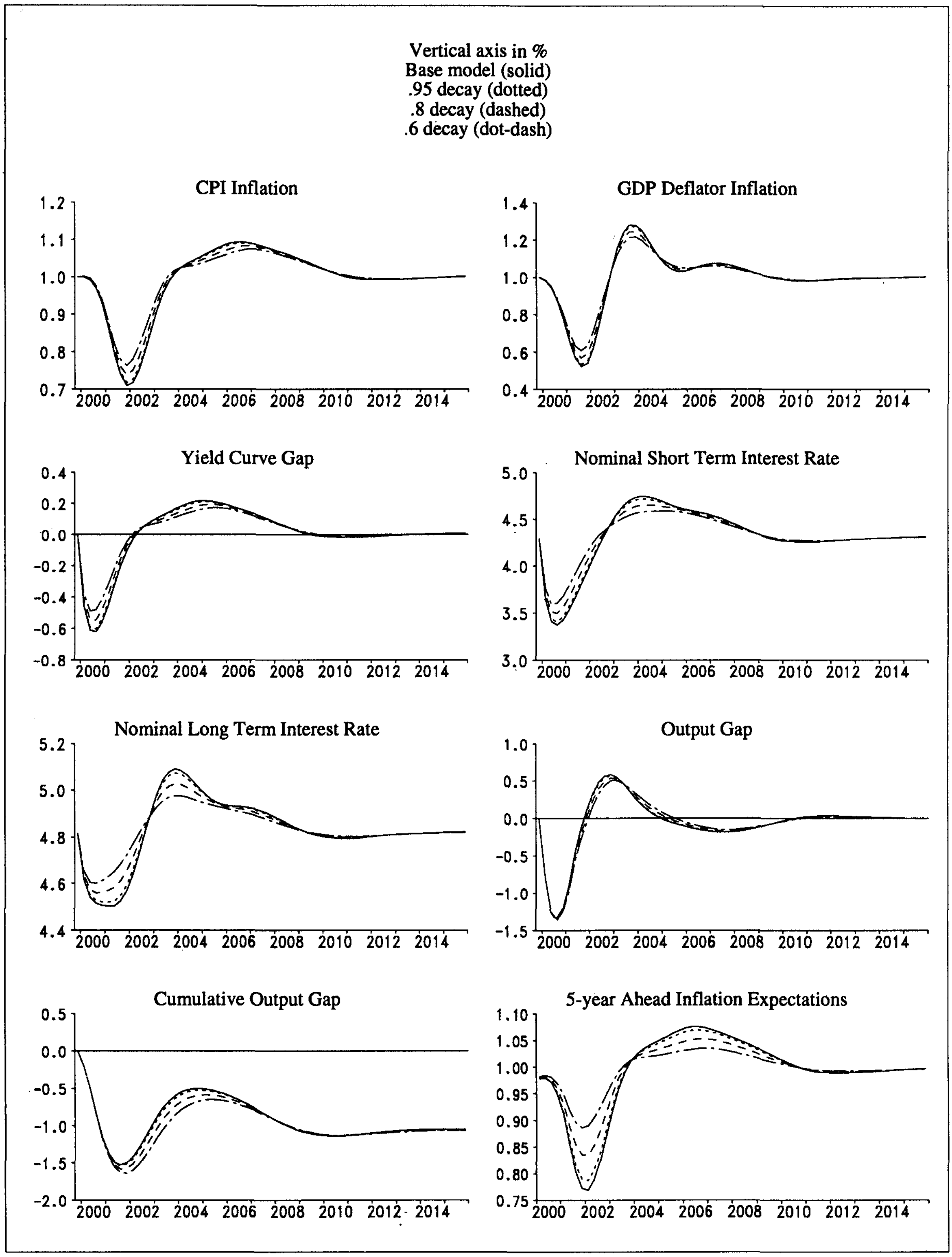


Figure 6

1% negative demand shock – fixed coefficients



on the target. However, the length of time it takes expectations to reach the new target is not greatly affected by the weight placed on the target. This is because the monetary authority, seeing a smaller gap between actual inflation and the target, makes less of an adjustment to interest rates. Consistent with this, the total length of time it takes for inflation to reach the new target is not altered significantly. The greater the weight on the target, the smaller the change in short-term interest rates in a disinflation shock, and thus the lower the cumulative output gap. With a 0.8 decay, nominal short-term interest rates peak at 5.3%, compared to 4.5% with a decay of 0.6.

In a negative demand shock, the greater the weight on the target, the more expectations act as an anchor. Inflation expectations do not exert as much downward pressure on inflation and therefore, it does not trough as low. With expectations acting as an anchor, the monetary authority does not have to respond as much. There is also less secondary cycling with greater weight on the target. While the cumulative output loss is greater in the early part of the demand shock, with greater weight on the target, the reduction in secondary cycling results in the total cumulative output loss remaining largely unchanged.

While these results are consistent with what is expected from an increase in credibility, the large impact which a change in the target can have on expectations is troubling, particularly since expectations will alter given an announced change in target, even if the monetary authority does nothing to implement its announced policy. This is illustrated in Figure 7, which shows a disinflation shock with a 1-year delay in the response of the monetary authority. In other words, the target rate of inflation was reduced by 1 percentage point, but the yield curve gap and nominal interest rates were set to remain unchanged until the end of four quarters.

It can be seen that even when the monetary authority does not follow a policy consistent with the target, inflation expectations fall immediately, causing in turn a fall in inflation. When the decay is 0.95, inflation falls only marginally (0.02 percentage points in the first quarter and 0.06 percentage points by the fourth quarter). With a greater weight, however, the effect is more pronounced. With a decay of 0.6, inflation falls 0.08 percentage points in the first quarter and 0.36 percentage points (or a third of the way to the new target) by the fourth quarter, before the monetary authority has begun to act. With a decay of only 0.8, the announcement effect is sufficiently strong that essentially no increase in short-term interest rates is necessary. In the base model, inflation and inflation expectations, do not adjust until the monetary authority starts to increase interest rates.

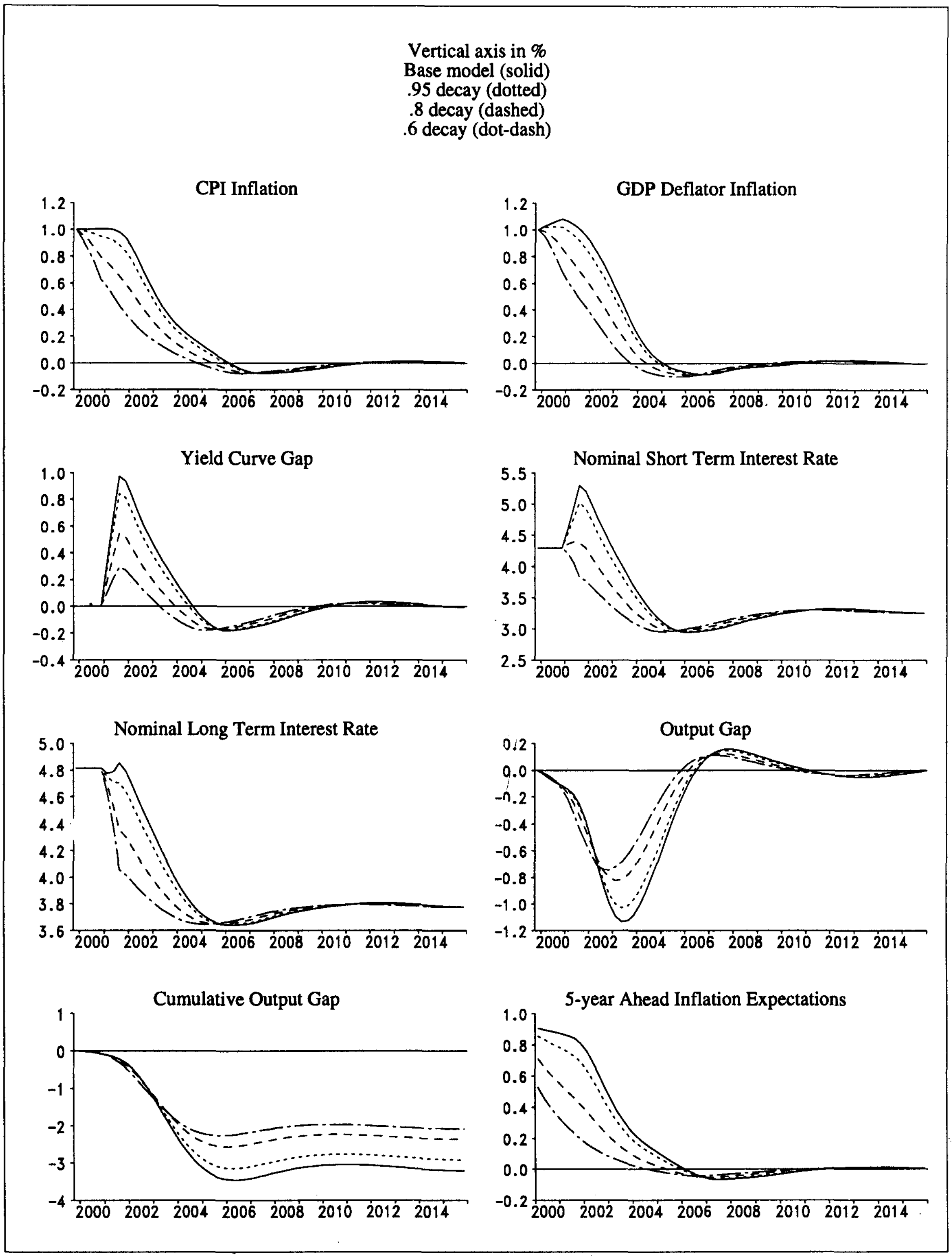
It was argued above that incorporating short-term expectations which are inconsistent with the underlying dynamics of the model is inappropriate. It is even less appropriate when expectations are highly inconsistent with the actions of the monetary authority. Incorporating excessive announcement effects could lead to an underestimation of the costs of reducing the target. As can be seen in these results, the pure announcement effect of a policy change can do much of the work towards bringing expectations in line with the new target, without any action needed on the part of the monetary authority. This decreases the onus on the monetary authority to try to adjust monetary conditions. The projection process, however, is one in which the Staff try to present a scenario where the risks are balanced. On such questions as the degree to which credibility will help the monetary authority the model should if anything err on the side of caution. (For similar reasons, steady-state QPM does not include real benefits from a lower rate of inflation, even though the Staff believe such benefits occur.) Incorporating the target directly into price expectations is not, therefore, a satisfactory way of introducing increased credibility effects into QPM.

3.2 Introducing the perceived target

An alternative to introducing the target directly into price expectations is to introduce a "perceived target", a variable which reflects what agents believe to be the actual target being used by

Figure 7

Delayed 1% disinflation – fixed coefficients



the monetary authority.[3] The perceived target is a mechanism through which expectations are sensitive to the action of the monetary authority but allow expectations to "look through" the short run price dynamics in the model.

The perceived target is calculated as a combination of a forward-looking component (model consistent inflation four and five years ahead) and a backward-looking component:

$$PCPITAR\_E=1.35*PCPITAR\_E(-1)-0.425*PCPITAR\_E(-2)+0.0375*(PCPI\_DA(16)+PCPI\_DA(20))$$

where *PCPITAR_E* is the perceived target inflation rate and *PCPI_DA(T)* is the model consistent inflation rate in period *T*.

The forward-looking component of the perceived target provides a measure of where people expect inflation will go over the medium to long term, once adjustment to current shocks has occurred. If the monetary authority is expected to keep inflation close to the target over the medium-term the model consistent inflation rate will be close to the target. If the monetary authority does not act in a manner consistent with its announced target, however, the perceived target may differ from the actual target.

Credibility can now be thought of as having two components. The perceived target is the policy which agents believe is being followed by the monetary authority, which may or may not be equivalent to the announced target. In other words, one measure of credibility is the difference between the perceived and announced target. The weight which the perceived target has within the expectations equation can be thought of as either the proportion of agents who use their perception of policy to form expectations, or the degree to which agents think policy will determine inflation outcomes.

This structure for expectations is very flexible for incorporating alternative assumptions about credibility. The perceived target can easily be replaced by other expectations processes, tying long-term expectations to real long-term bond rates, for example. The coefficient on the perceived target could also be made endogenous, providing another means of gaining or losing credibility over time.

Incorporating a perceived target also provides considerable flexibility for calibrating different effects associated with credibility. This can be seen by reviewing changes to model properties given changes to the values of the key components within the perceived target framework: the choice of time horizon for the forward-looking component, the weights on the forward and backward components within the perceived target, and the overall weight on the perceived target within price expectations.

Time horizon for the forward-looking component

The choice of time horizon for the forward-looking component is based on the idea that it reflects people's medium-term inflation expectations. The horizon 4-5 years ahead was selected because it is far enough ahead not to be greatly affected by the short-term effects of shocks to the economy. This can be seen in Figure 8, which shows a demand shock with time horizons for the forward-looking component of 2-3 years, 3-4 years and 4-5 years. A high weight has been given to the perceived target in price expectations in these simulations (0.56) so as to clearly illustrate the effects of varying the time horizon. The perceived target shows much greater variation when a shorter time horizon is used. It should be noted that in a negative demand shock, the perceived target increases rather than decreases, since it is picking up the secondary cycling which occurs after 2002. The 4-5

[3] This is the approach used in CPAM, where a "perceived" target is included in inflation expectations. Richard Black, Tiff Macklem and David Rose, "On Policy Rules for Price Stability", Bank of Canada Conference 3rd-4th May, forthcoming, p. 13.

Figure 8

1% demand shock – varying time horizon for forward-looking component

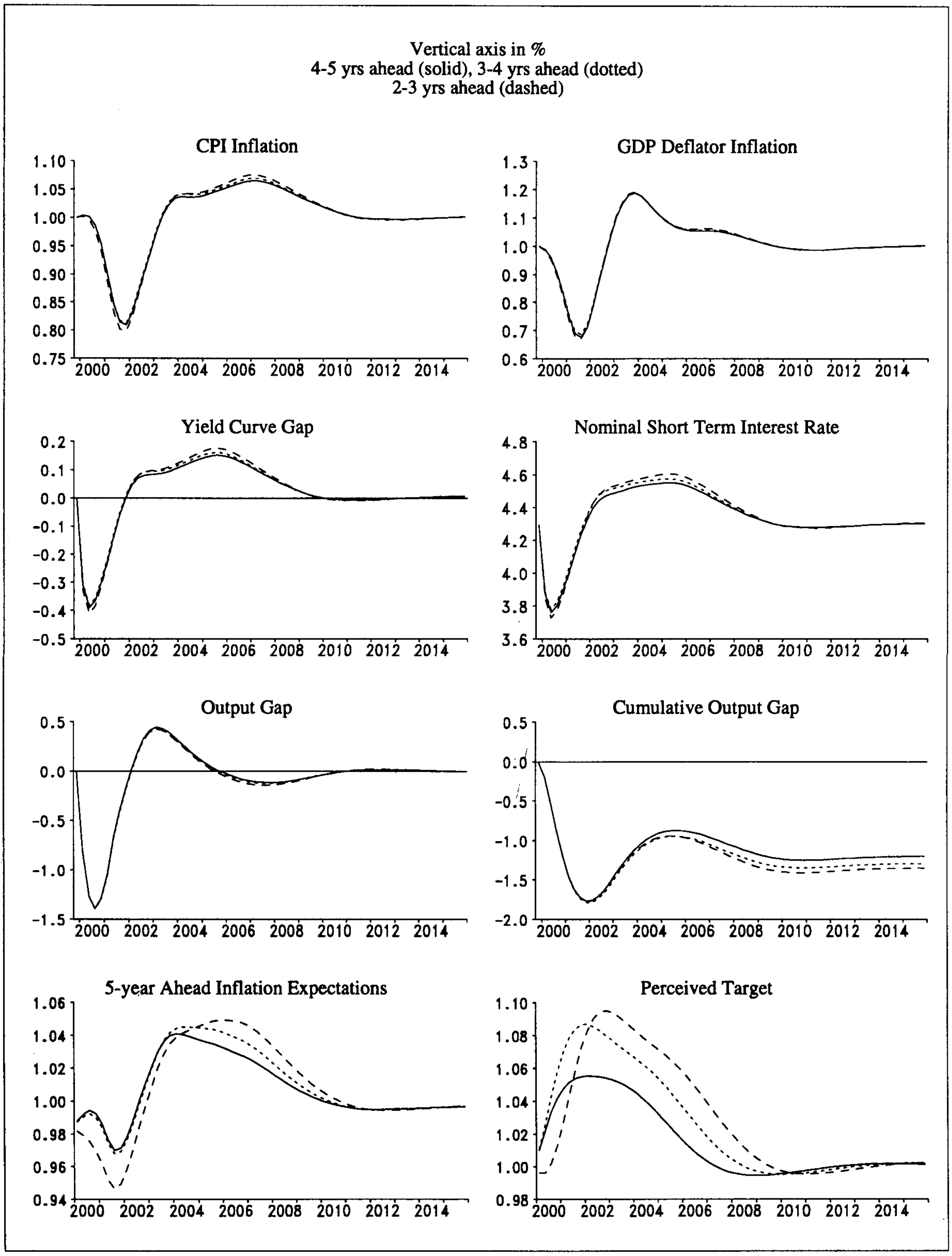
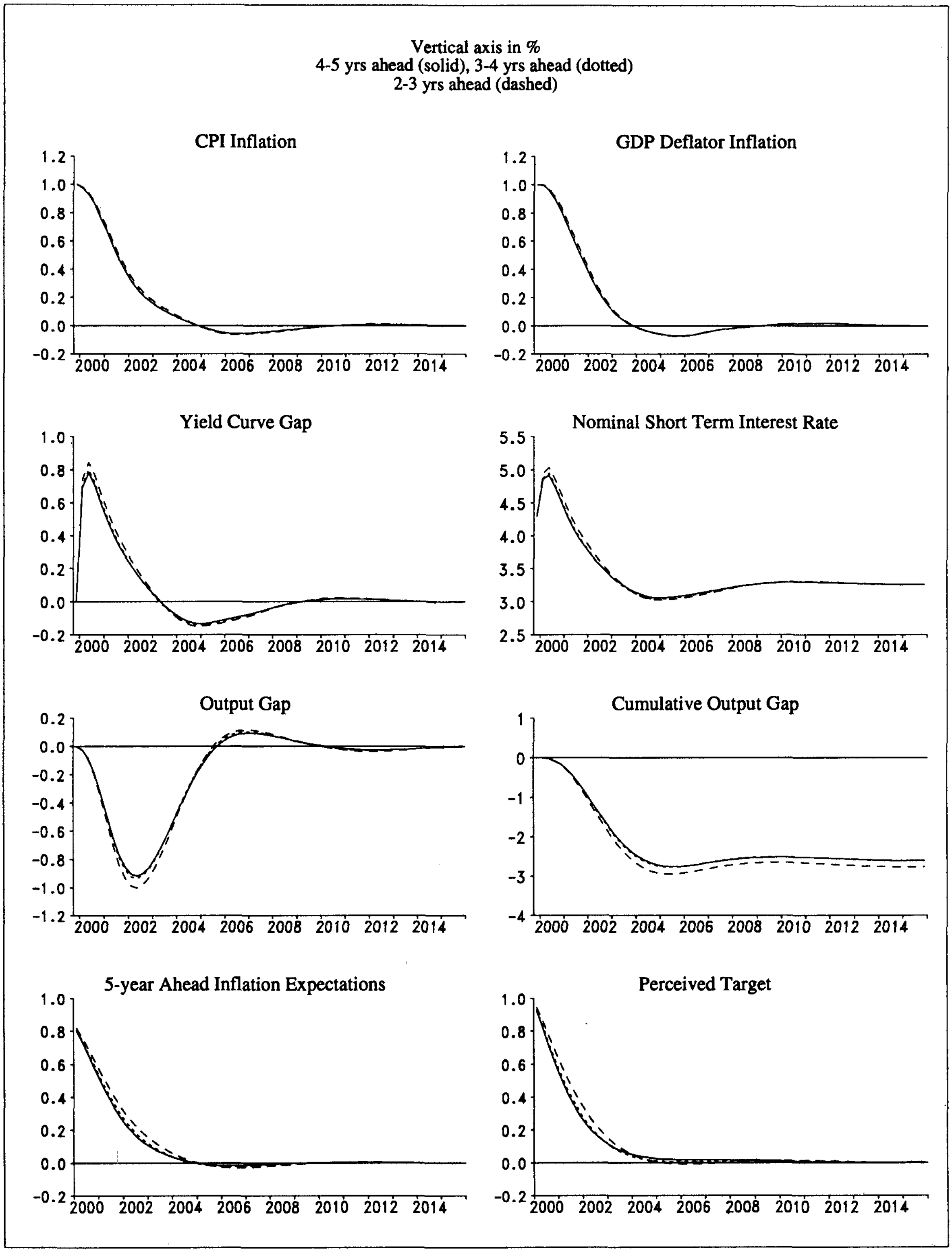


Figure 9

1% disinflation – varying time horizon for forward-looking component



year time horizon shows relatively little variation since it looks beyond the effects of the shock. Clearly, it would not be appropriate to have a large positive increase in the perceived target during a negative demand shock, thus the 4-5 year time horizon is the most appropriate. The path of the output gap is similar in all three cases. The cumulative output gap is slightly smaller with the 4-5 year time horizon, however, due to less secondary cycling.

Figure 9 shows a disinflation shock for the same time horizons. In the disinflation shock the 3-4 and 4-5 year horizons give very similar results since both are looking ahead to the new target almost immediately. The 2-3 year time horizon, however, causes inflation expectations to adjust more slowly. For this reason, the policy response in the 2-3 year scenario is greater, resulting in a larger sacrifice ratio.

Weighting of components within perceived target

Credibility is lost initially if the forward-looking component varies from the actual target. This may be due to the actions of the monetary authority, or it may be due to the impact of shocks to the economy if the time horizon for the perceived target is too short. The coefficients on the two lags of the perceived target then determine the speed with which a change in the forward-looking component leads to further loss of credibility. The greater the coefficient on the first lag, the faster the change in the perceived target. For CPAM the coefficients were chosen so that the roots of the equation are 0.85 and 0.50.[4]

The overall weighting on the backward as compared to forward elements of the perceived target has no impact on shocks to the economy where the target does not change. (This result assumes, of course, that the time horizon chosen for the forward component looks beyond the shock.) The relative weighting on backward and forward elements does, however, have a large impact on shocks where the target changes, since it determines the speed with which a change in the target is reflected in the perceived targets, and thus the monetary response and cumulative output gap in a disinflation or inflation shock.

The effects of changing these weights can be seen in Figure 10 which shows a disinflation shock for different weights on the backward and forward components. Again, a relatively large weight (0.56) was chosen for the perceived target within expectations, to illustrate the impact of the coefficient changes. Increasing the weight on the forward-looking component within the perceived target increases the speed with which it reaches the new target. With a weight of 0.075 on the forward component, the perceived target falls around two-fifths of the way to the new target by the end of the first year. By the end of the third year it has almost reached the target, but does not fully converge until the fifth year. By contrast, with a weight of 0.445 on the forward component the perceived target has reached the actual target by the end of the first year.

Consistent with the change in the perceived target, the greater the weight on the forward component, the faster inflation falls, and the less the monetary authority has to alter interest rates. With a weight of 0.445 on the forward component, short-term nominal interest rates do not increase at all. The effect of increases in the forward-looking component on the cumulative output gap is highly non linear and decreases as the weights increase i.e. increasing the weight from 0.075 to 0.1675 causes the cumulative output gap to fall to 2.2% as compared to about 3% when the weights are increased from 0.168 to 0.445.

The implications of different weights within the perceived target are even more obvious if real data are used. Figure 11 shows the values of the perceived target, along with actual year-on-year inflation, calculated over history. There is a considerable difference between the different

4 The coefficient on the first term is the sum of the two roots, whereas the coefficient on the second lag is the product of the two roots.

Figure 10

1% disinflation – varying backward for forward-looking components

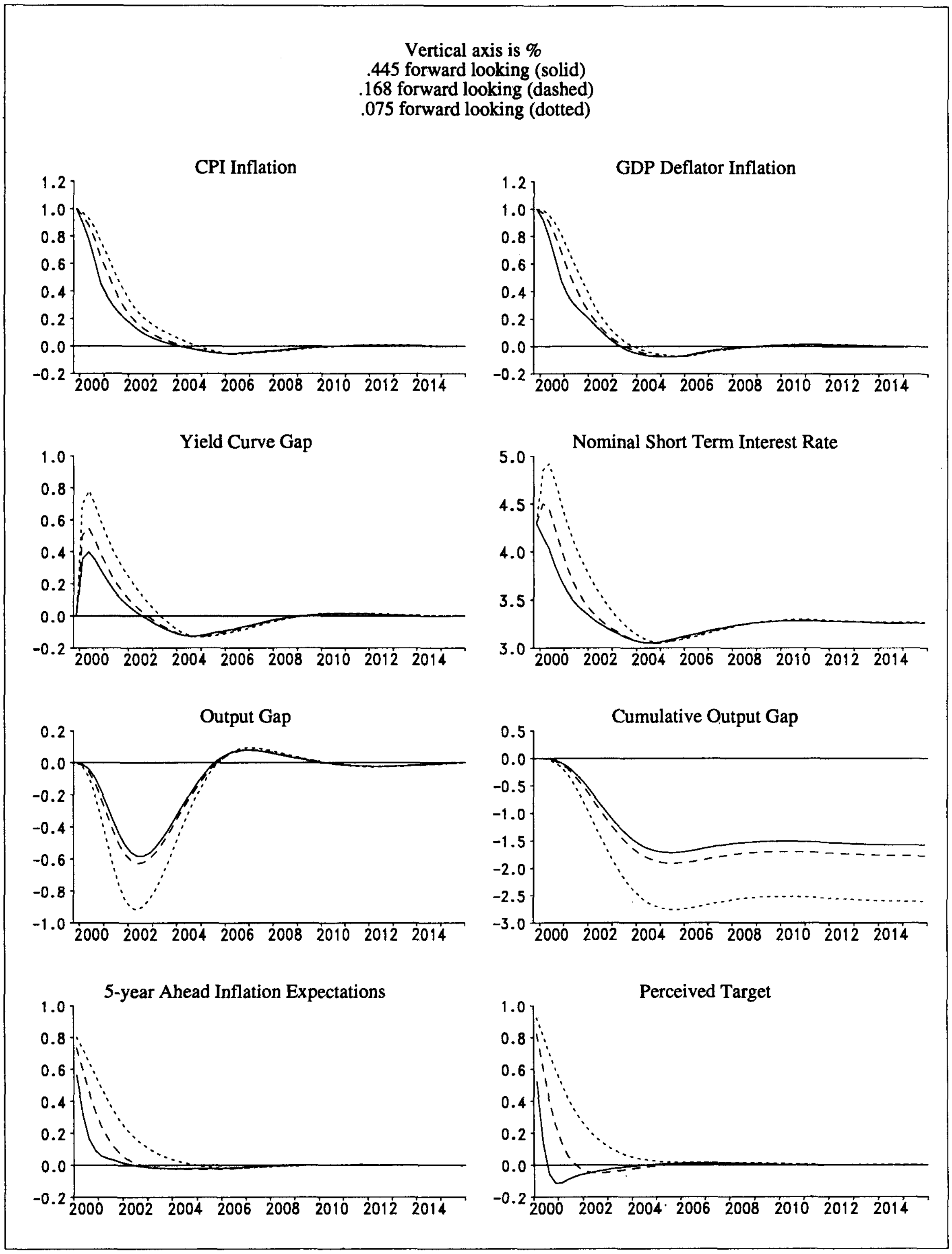
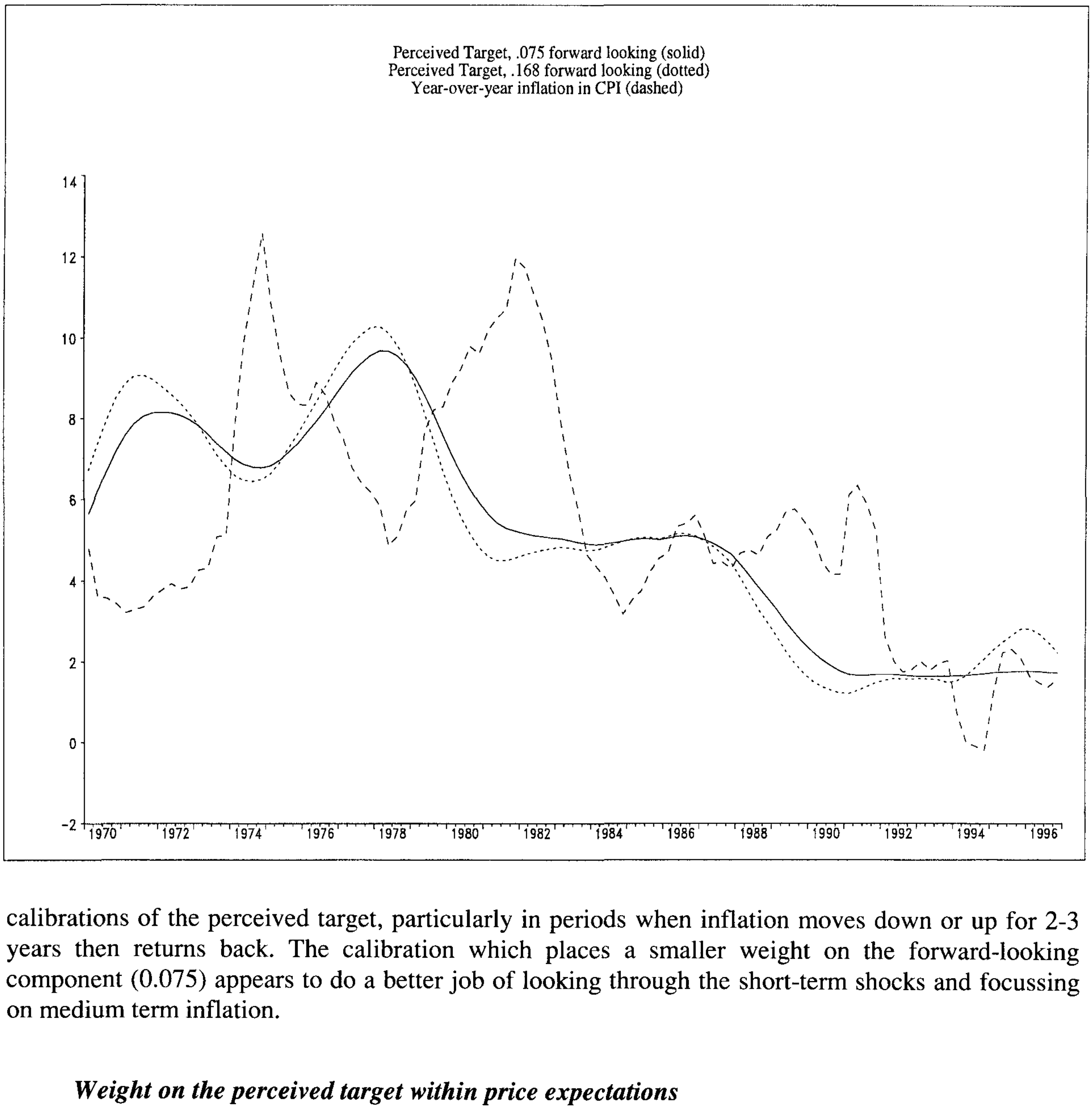


Figure 11

The perceived target over history



calibrations of the perceived target, particularly in periods when inflation moves down or up for 2-3 years then returns back. The calibration which places a smaller weight on the forward-looking component (0.075) appears to do a better job of looking through the short-term shocks and focussing on medium term inflation.

Weight on the perceived target within price expectations

While the relative weights on the forward and backward components of the perceived target affect policy shocks, the weight on the perceived target within price expectations affects all shocks. Figures 12 and 13 show the impact of varying the weight on the perceived target in the disinflation and demand shocks. These results are very similar to those resulting from incorporating the target directly into price expectations. In a disinflation shock, expectations adjust more quickly and inflation falls more in the first year. Inflation also locks into the target more neatly, without as much secondary cycling. The overall time it takes to reach the new target is not, however, greatly changed. The degree to which the monetary authority needs to adjust interest rates decreases the greater the weight on the perceived target, particularly in the first year, and the cumulative output gap falls. In the demand shock, the variability of expectations and inflation falls, with both a shallower trough and less secondary cycling. Again, the monetary authority needs to do less to offset the shock, and thus the cumulative output gap increases.

Incorporating the perceived target reduces, though does not entirely eliminate, the problem of a "pure" announcement effect where expectations adjust even when the monetary

Figure 12

Disinflation, the perceived target – varying the weight within expectations

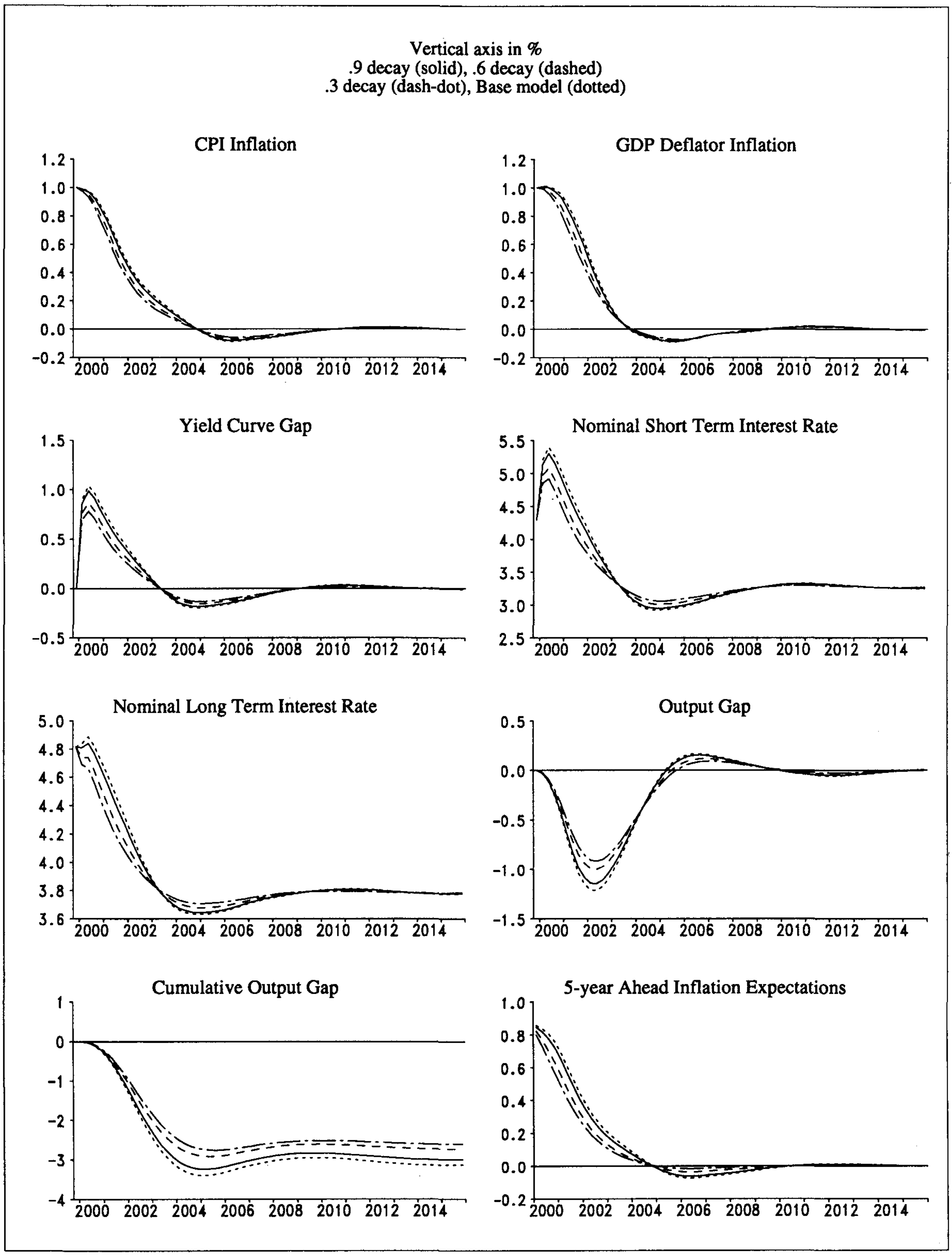


Figure 13

Demand, the perceived target – varying the weight within expectations

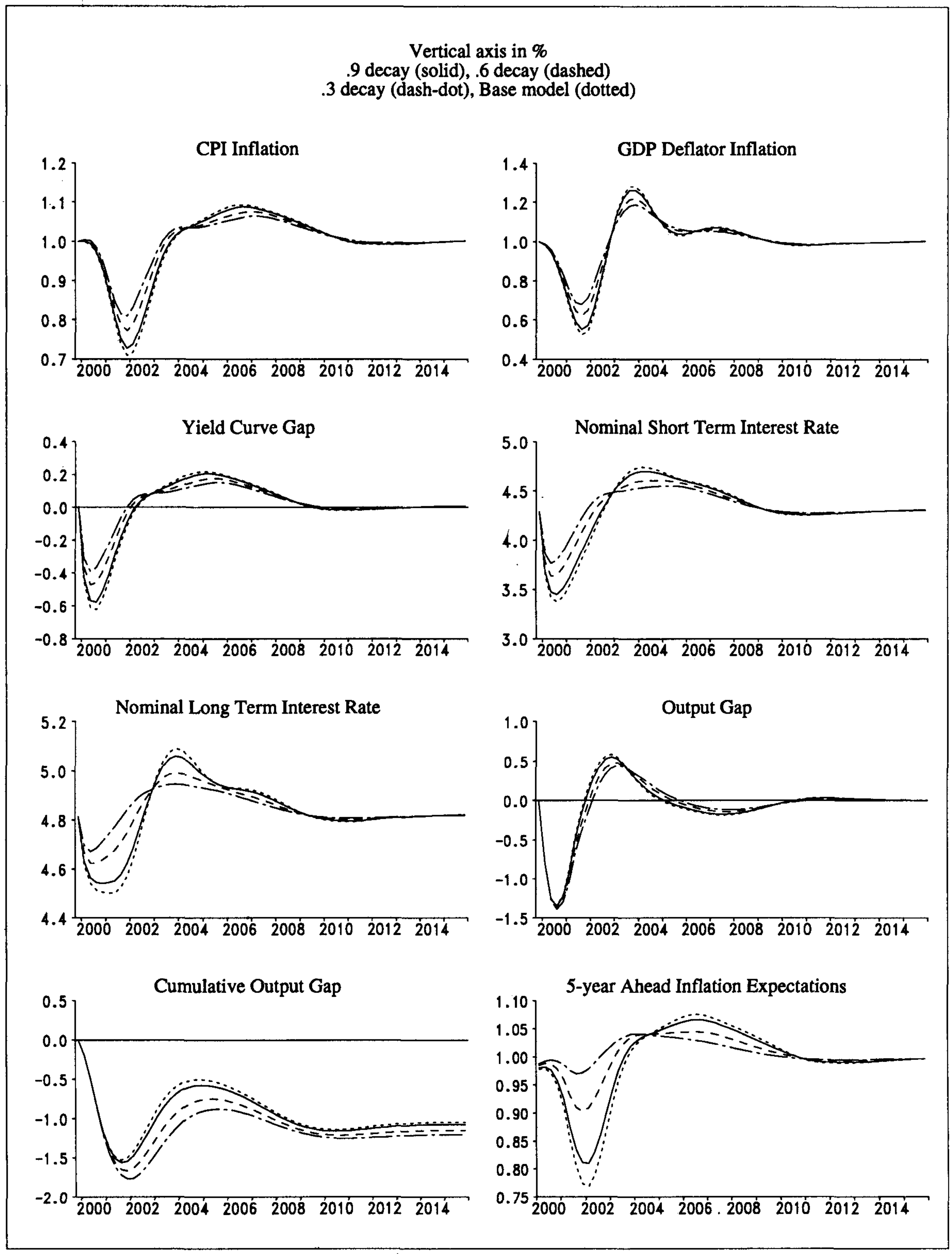
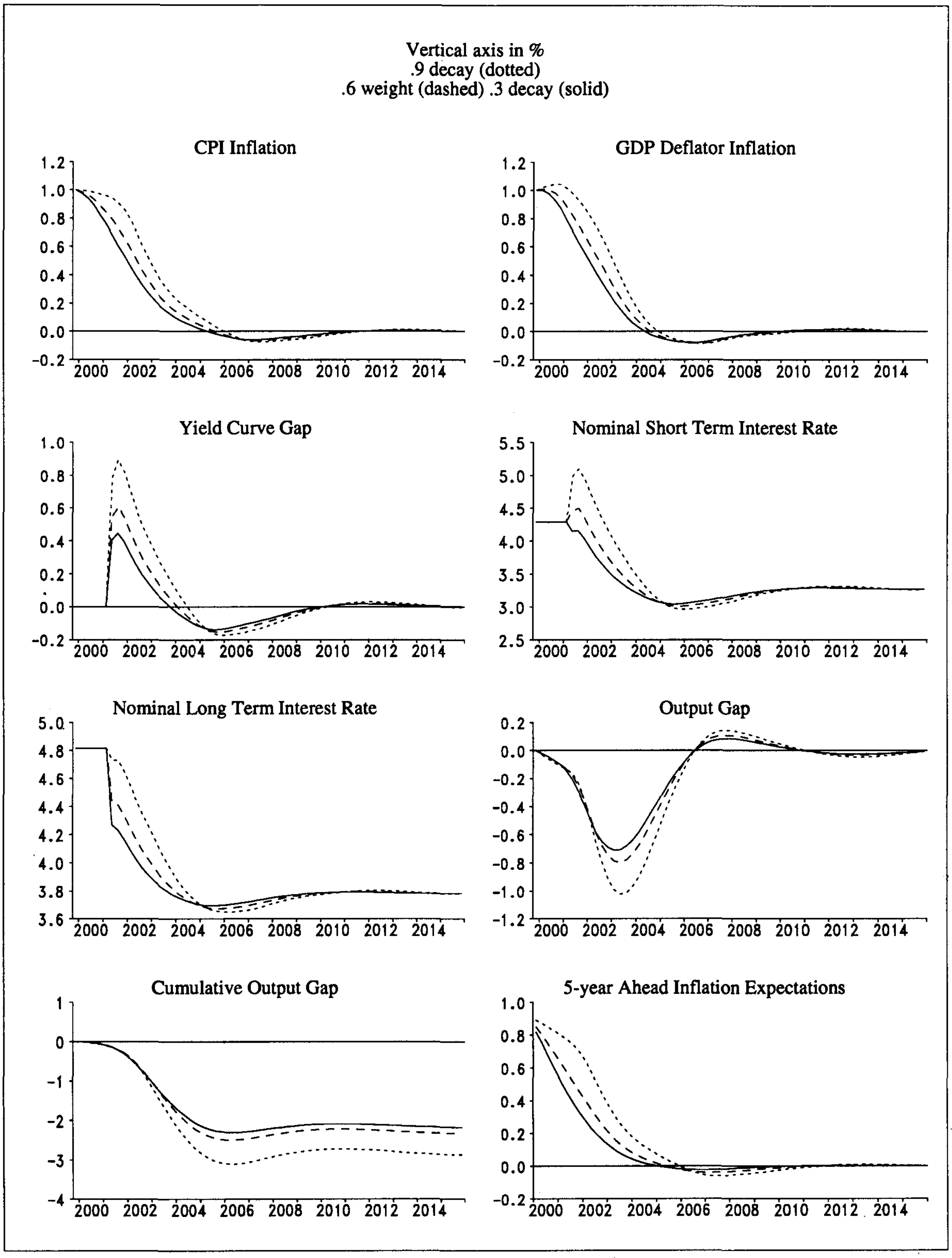


Figure 14

Delayed response disinflation – with perceived target



authority is not acting consistently with the stated target. This is illustrated in Figure 14, which shows the impact of a disinflation shock where the monetary authority does not react for one year. When the perceived target has a low weight, expectations and inflation do not decline until the monetary authority starts to increase interest rates. When the perceived target has a high weight, however, there is the potential for an announcement effect. This is because given the reaction function in QPM, once the monetary authority starts to act, it does so very fast and still brings inflation down to the target by the 4-5 year time horizon being used in the forward-looking component. The longer the period over which the monetary authority fails to act, however, the less likely that expectations will adjust. If the monetary authority fails to act for four years, for example, there is no announcement effect with the perceived target. With the actual target in expectations, there would still be an announcement effect in such a case.

While building in an effect which forces expectations to move towards the target independent of the actions of monetary policy, is not appropriate, it must be remembered that incorporating an announcement effect is an implication of credibility. Increased credibility implies that the announcement of a new policy will have a greater immediate impact on expectations. The size of the announcement effect is a calibration issue.

The above analysis shows, therefore, that the perceived target framework offers considerable flexibility in incorporating credibility effects in QPM. Consistent with what one would expect from increased credibility, the perceived target acts as an anchor for price expectations in non-policy shocks, and increases the speed with which expectations adjust when the target adjusts. In both kinds of shock, this reduces the extent to which the monetary authority must adjust interest rates. The greater the weight on the perceived target within expectations, the greater the impact it has on both kinds of shocks. The relative impact on policy and non-policy shocks can be altered by varying the weights on the backward and forward-looking components within the perceived target.

4. Calibration of the model

Having decided on a framework for incorporating a specific credibility effect within QPM, it is necessary to calibrate the equations. In other words, weights must be chosen so that the dynamics of expectations reflect empirical evidence on the formation of expectations in recent years. Values must also be chosen for other key elements which are calibrated such as the sacrifice and benefit ratios. This section provides an overview of empirical evidence on the effects of credibility and some preliminary empirical work to establish a set of stylized facts to be used for the calibration. An initial calibration of the model is chosen and the implications are assessed for changes in model properties. Since the data on expectations and the effects of credibility are sparse, there are perhaps more questions raised than substantive answers given. These provide a useful guide, however, for future research.

4.1 Empirical evidence of increased credibility

Despite widespread acceptance of the benefits of credibility and adoption in many countries of low-inflation policies designed to build credibility, empirical evidence about credibility is very inconclusive.[5] There are two main ways in which people have tested for changes in credibility: analysing changes in the inflation/unemployment or inflation/output trade-off, and identifying and modelling changes in inflation expectations. The literature on sacrifice ratios has uncovered little evidence of increased credibility, while studies focussing more directly on inflation expectations have produced some results which are consistent with growing credibility.

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5 For reviews of the empirical literature see Blackburn and Christensen (1989), Amano et al. (1996) and St. Amant (1997).

Blackburn and Christensen (1989) start with the hypothesis that increased credibility will reduce the inflation/unemployment trade-off. They calculate these trade-offs for disinflations in three countries, all of which made a commitment to some nominal target and many announcements about new anti-inflation policy designed to influence expectations. They found no declines in the inflation/unemployment trade-off, suggesting that the policies did not affect expectations. Debelle (1996) finds no reduction in the inflation/output trade-off in the recent disinflation in Australia, New Zealand and Canada. He finds, for example, that the sacrifice ratio in Canada was the highest ever in the last disinflation.

Changes in sacrifice ratios over time may be a misleading indicator of monetary policy credibility. In Canada, for example, it can be argued that the high sacrifice ratio in the early 1990s, was in part related to good credibility: inflation was generally below the mid-point of the target bands thus if policies were credible, people would have been expecting inflation to rise back up to the mid-point. The disinflation also came at a time of considerable restructuring in the economy, which may have contributed to short-term declines in output.

There is also some doubt as to what implications increased credibility in a low inflation environment will have for the real economy. A number of models of price behaviour imply non-linearities in the Phillips Curve which, by decreasing the coefficient on the output gap, would tend to increase output/inflation trade-offs in regimes with a lower level and/or less volatile inflation, unless there was a significant change in the process generating inflation expectations (for example, through increased credibility). Signal extraction models, for example, suggest that the coefficient on the output gap is lower when inflation volatility is lower, because agents are better able to distinguish between relative price and aggregate price shocks. Adjustment cost models suggest a similar relationship between the coefficient on the output gap and the level of inflation. In low inflation regimes, for example, agents may negotiate longer contracts in order to save on adjustment costs, which will would tend to slow adjustment. Dupasquier and Ricketts (1997a, b) test for a variety of such non-linearities in the Phillips curve. For Canada they find evidence of non-linearities. While the strongest evidence appears to support non-linearities with respect to the output gap, they are unable to rule out effects coming from the level and the volatility of inflation. To the degree that the volatility of inflation falls when credibility is increased, therefore, there may be an offsetting impact on the beneficial influence of credibility on the sacrifice ratio.

An alternative way of trying to gauge the degree of credibility is to determine whether survey data on inflation forecasts change over time in a manner consistent with the policy objective. With increased credibility, survey expectations will likely show less variability since they will be more anchored to the target, and the dispersion of individual forecasts will be lower. Expectations should also adjust more quickly to changes in policy. Debelle (1996), for example, models inflation expectations as an autoregressive process and searches for evidence of a structural break in recent years. He finds evidence consistent with increased credibility for New Zealand, no evidence of increased credibility for Australia, and mixed evidence in Canada. Bachelor and Orr (1991) and Fischer and Orr (1994) look at measures of uncertainty based on the variance of inflation expectations across forecasters for the United Kingdom and New Zealand. While their results show declining uncertainty, this seems largely to be due to lower rates of inflation.

Johnson (1997) uses survey data from professional forecasts across 18 countries for the period 1984 to 1995. He finds that in most cases, in both inflation targeting and non-targeting countries alike, the disinflations of the 1990s were unanticipated which led him to conclude that inflation targets were not instantly credible. He does find evidence of decreased variance of forecast errors in recent years, but is unable to differentiate between the effects of targets and the effect of a period of more stable inflation. Perrier (1997) applies and extends Johnson's methodology to data on inflation expectations in Canada from the Conference Board Survey of Forecasters. He finds that explicit targets contributed to reducing the mean and variance of forecast errors, and concludes from this that the targets did increase credibility.

Work on inflation expectations and regime switching models also provides evidence that the inflation process has changed, and that these changes are consistent with improvements in credibility.[6] Fillion and Léonard (1997), for example, estimate a Phillips Curve for Canada using inflation expectations based on a Markov switching model estimated by Ricketts and Rose (1995). In the model there are three possible inflation regimes: low and stable inflation, moderate inflation and very high (unit root) inflation. The model suggests that since 1991 a very high probability is placed on being in a low inflation regime (over 90%). This supports the idea that expectations are in line with the current target, but it does not give clear indications about changes in credibility over time. In the second half of the 1980s a similar probability was attached to being in a moderate inflation regime with a mean inflation rate of around 4%. Similarly, the transitions from one regime to another do not suggest expectations are adjusting more quickly to policy changes than in the past. On average transitions take about 2 years. The transition to the low inflation regime of 1990s took only 6 quarters, but the fastest transition was the adjustment to a high inflation regime in 1974.

Overall, therefore, while the empirical evidence based on expectations is consistent with increased credibility it does not provide any quantitative benchmarks which can be used in model calibration. Similarly, empirical work provides no evidence on changes in sacrifice ratios. This is not surprising given that the use of low inflation policies specifically designed to build credibility is relatively recent. It is quite possible that the period has not yet been long enough to build significantly greater credibility or for credibility effects to show up in the data. It nevertheless presents a problem for modellers trying to incorporate credibility effects.

4.2 Overview of measures of expectations for Canada

Survey-based expectations

The main source of survey-based inflation expectations in Canada is the Conference Board Survey of Forecasters. The survey covers from 8 to 17 private sector forecasters and collects forecasts of annual average inflation in the current year and one year ahead, for both the CPI and the GDP deflator. It is available on a quarterly basis from 1984 and on an annual basis from 1975.

Figure 15 shows consensus (mean) one-year-ahead inflation expectations of the CPI from the Conference Board Survey, compared to actual annual average inflation at the time the forecast was made. It is evident that these forecasts closely track contemporaneous inflation.[7] In the disinflations of the early 1980s and 1990s, expectations were generally below actual inflation, but above realised inflation. In the 1990s, expectations have also closely followed the current mid-point of the target range and have been around 2% since 1993.

Consensus Economics also provides quarterly forecasts of inflation. They survey a very similar group to the Conference Board, but provide forecasts over a greater range of time horizons, up to ten years ahead.[8] Unfortunately, the data only begin in 1991. Figure 16 shows CPI inflation forecasts one, two, five and ten years ahead. Not surprisingly, given the similar pool of forecasters used, the one-year-ahead forecast is almost identical to that from the Conference Board. These data suggest that longer term inflation expectations adjust at a similar speed as shorter term expectations. Despite their very different time horizons, the two-year, five-year and ten-year ahead forecasts show a very similar evolution over the 1990s, again closely related to the contemporaneous mid-point of the target range and to the contemporaneous rate of inflation.

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6 In particular see Laxton, Ricketts and Rose (1993), Ricketts and Rose (1995) and Fillion and Léonard (1997).

7 The fall in annual average inflation in 1994 is largely associated with a decline in the indirect tax on tobacco, which was unanticipated.

8 Forecasts are published in *Consensus Forecasts*. Longer term horizons are available semi-annually.

Figure 15

One-year-ahead inflation expectations, Conference Board Survey of Forecasters

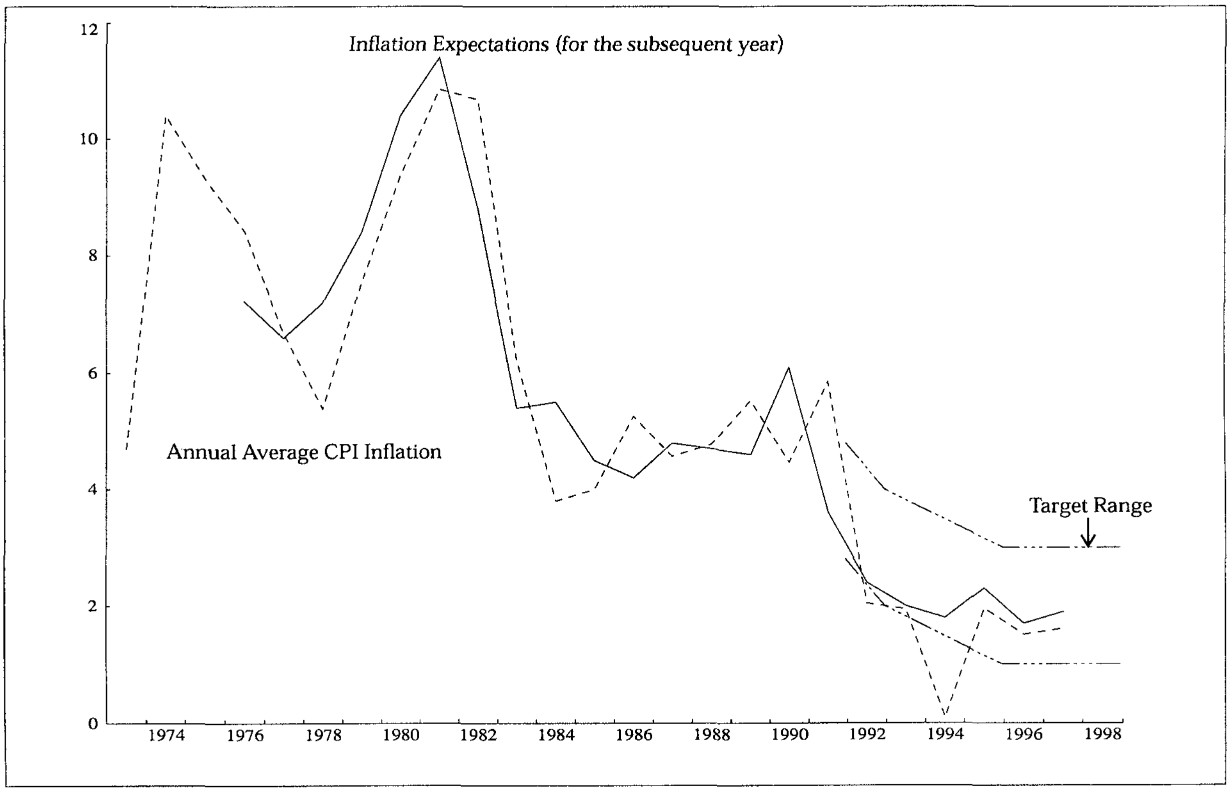
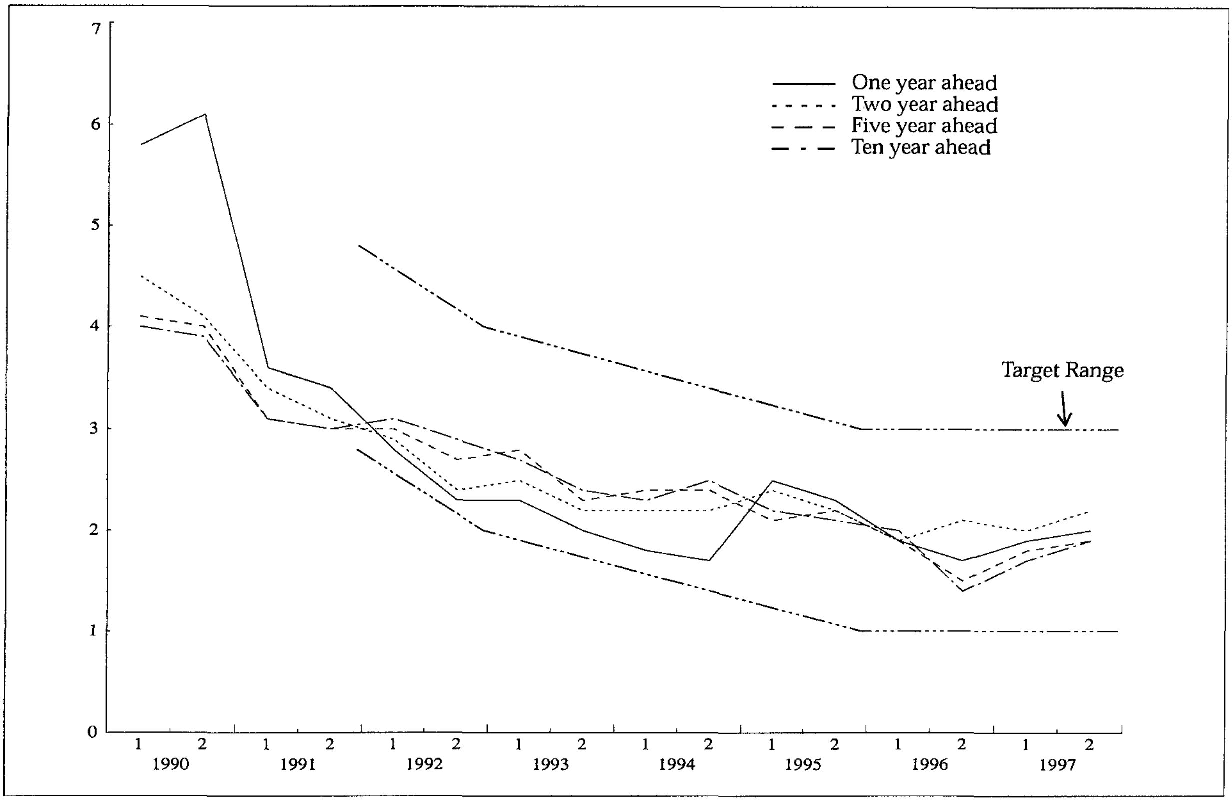


Figure 16

Consensus forecasts of CPI inflation (plotted against date forecast was made)



Given that the first announcement of a path for inflation targets was made in February 1991, in theory the adjustment of expectations in 1991 compared to 1990 should reveal something about the announcement effect, perhaps provide a lower bound since credibility was likely to be lower rather than higher than currently. In practice, however, it is hard to gain much information from the evolution of expectations over this period. The one-year ahead expectations held in 1990 were greatly affected by the pre-announced introduction of GST for 1991, thus a large decline in expected inflation would be expected even in the absence of inflation targets. (On the positive side, though, it is clear that people did expect GST to have only a one-off impact on inflation.) Two-year ahead expectations fell a full percentage point from 4.1% in the second half of 1990 to 3.1% by the second half of 1991. This could be taken to suggest a very substantial announcement effect. Alternatively, it could be seen as following the path of underlying inflation. Moreover, the five and ten-year ahead expectations adjusted slightly more slowly than the two year-ahead expectations, which is less consistent with high credibility.

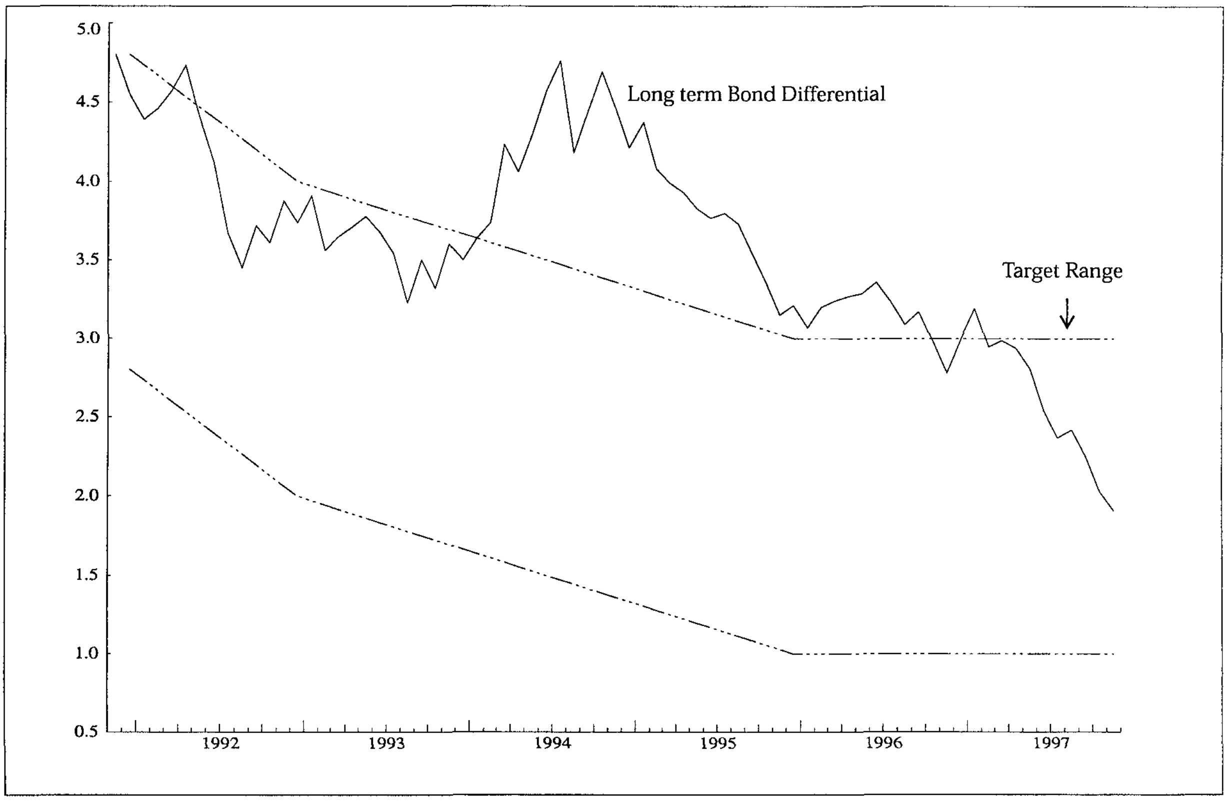
Real return bonds

Another source of data on long-term inflation expectations is the differential between conventional and real return bond yields. Real return bonds were first issued in December 1991. They are 30-year bonds, for which both the coupon payments and the principal are indexed to the CPI. The difference between the real return yields and the yield from a conventional bond of the same maturity is strongly associated with inflation expectations.

The differential between the two may not be identical to inflation expectations, since there are a number of things which may cause people to want different real returns on different

Figure 17

Differential between nominal and real long-term bonds



bonds.[9] For example, currently the secondary market for real return bonds is much smaller than that for conventional bonds, thus the real return bond is not as liquid. Agents may demand a premium on the real return bond as compensation for the greater liquidity risk, in terms of a higher real return relative to that expected from conventional bonds. If this is the case, the differential between the two will underestimate inflation expectations. Similarly, if there is considerable uncertainty over the future rate of inflation, investors in conventional bonds may demand a premium over and above that needed to compensate them for the average rate of inflation. In this case the differential between the two would overestimate inflation. Moreover, moving towards a situation of reduced uncertainty about inflation would cause a greater fall in the differential than can be accounted for by changes in inflation expectations.

Figure 17 shows the differential between the real return and conventional bonds, along with the announced path of inflation targets. Since 1992 the differential has fallen from over 4% to under 2%. The differential increased in 1994, but this was related to concerns over fiscal policy. Over this period long-term nominal rates increased in the United States, and rates in other industrialized countries followed suit. In those countries with high indebtedness, including Canada, greater increases occurred than elsewhere.

As mentioned above, the level of the differential is not necessarily a good representation of the level of inflation expectations, but the change in the differential is certainly consistent with declining expectations. Moreover, expectations have on average declined at a very similar speed to the actual path of announced targets, but with much greater variance than the Consensus Economics forecast. As before this evidence is not conclusive, particularly since inflation uncertainty likely declined over this period. Nevertheless, it provides another source of information on inflation expectations which is clearly suggestive that the targets have considerable credibility.

4.3 Establishing a benchmark calibration of QPM

Given the lack of conclusive empirical evidence, a cautious approach was taken in calibrating the credibility effect into QPM. Clearly, incorporating a large effect, which would imply a very large change in model properties such as the degree of monetary response needed in the face of shocks or policy changes and the resulting output gains and losses, cannot be justified. The weights chosen for the perceived target were those which provided the smoothest path for the perceived target over history: a weight of 0.075 on the forward-looking component, and of 1.35 and -0.425 on the first and second lags of the backward component (as discussed in Section 3.2). This path most closely reflects the idea that the perceived target reflects agents' medium-term inflation expectations. A decay of 0.8 was chosen, which gives a weight of 0.16 on the forward-looking component of the CPI.

As shown above, changing expectations in this manner would, all else equal, reduce the sacrifice and benefit ratios in a disinflation and an inflation. It is not clear, however, whether such an adjustment is appropriate. As described in Section 4.1, particularly in the discussion of Dupasquier and Ricketts (1997a, b), there are questions as to whether offsetting changes in the relationship between output and inflation may partly or fully counteract the gains from the faster adjustment of expectations, perhaps leading to no significant changes in sacrifice and benefit ratios associated with policy changes.[10] If this is the case, it may not be appropriate to let these ratios adjust in the absence of a new benchmark established on the basis of empirical work, rather than model simulations.

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9 For more information see Coté et al. (1996).

10 It is important to distinguish between two experiments i) where only credibility changes and ii) where both credibility and the level of inflation have changed.

Figure 18

Disinflation shock – with perceived target

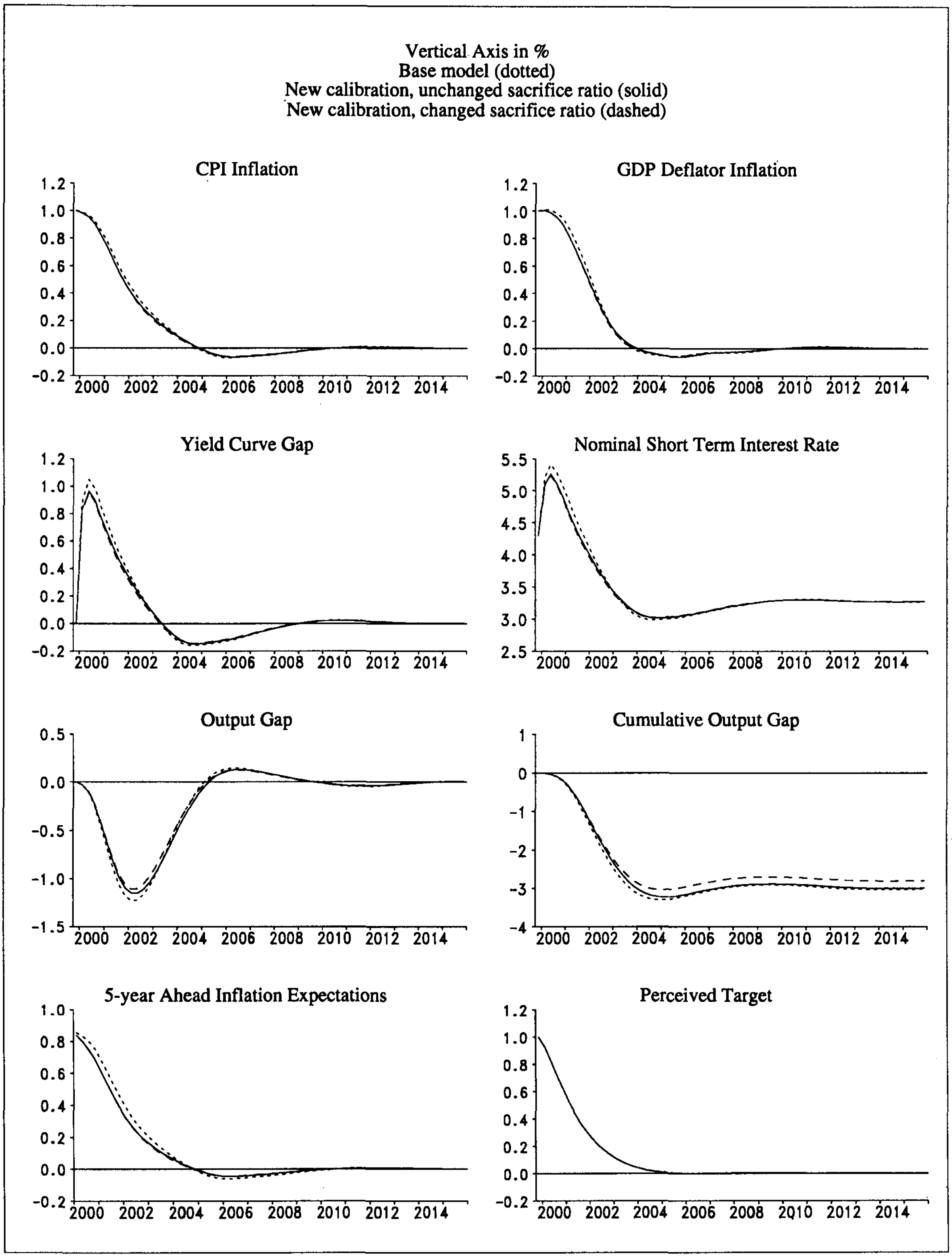
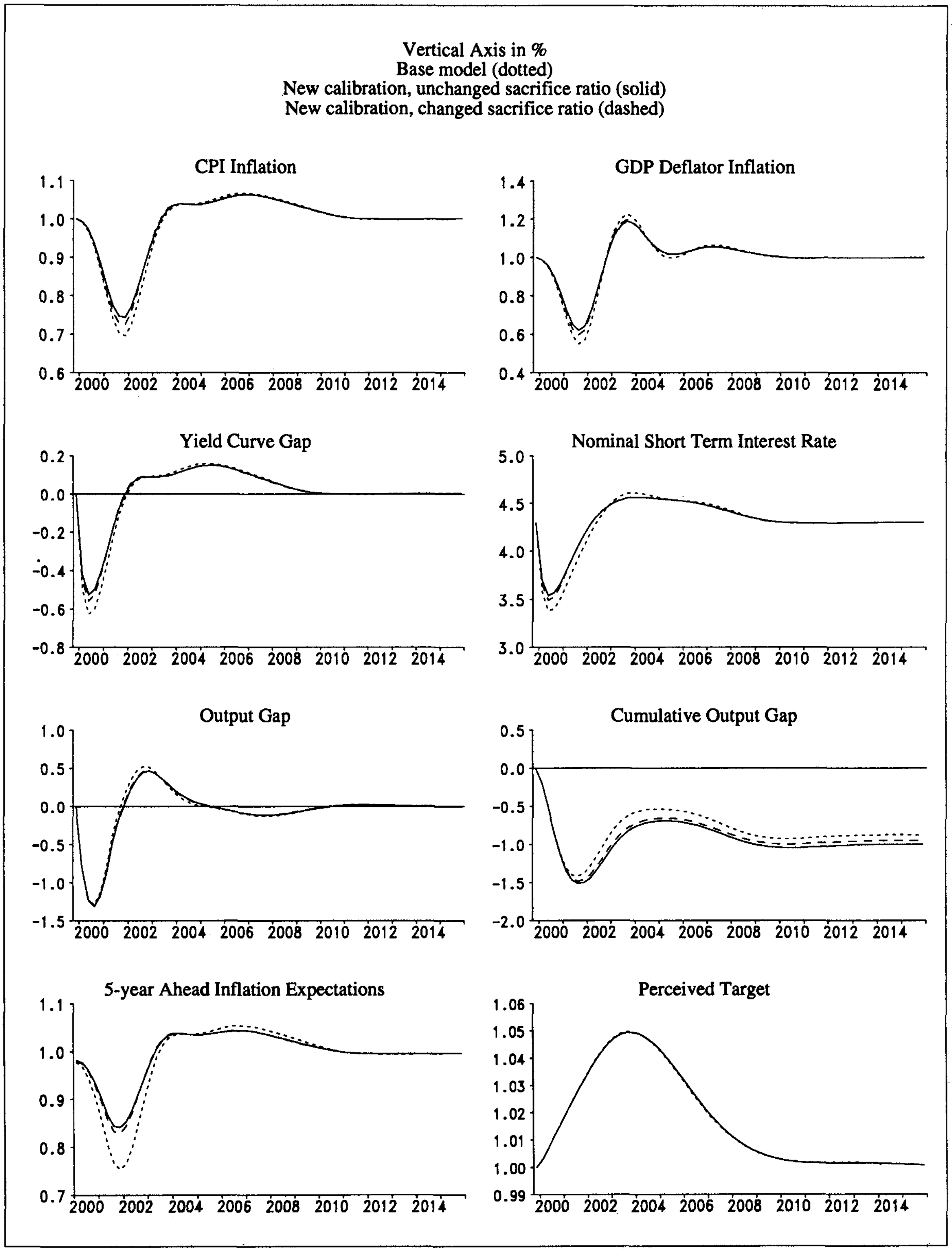


Figure 19

Demand shock – with perceived target



Given this uncertainty, two calibrations of the model are shown below: one in which the sacrifice and benefit ratios associated with a disinflation and inflation respectively, are restored to those in the base model;[11] and one where they are allowed to fully adjust with the introduction of the perceived target.

Figure 18 shows the implications of these two calibrations for a disinflation. Both models which include the perceived target give very similar results. The predominant change is the initial decline in inflation expectations compared to the base model, which reduces the initial monetary response needed to bring about the disinflation. In the model where the sacrifice ratio is allowed to adjust, monetary policy eases a little more quickly in the second and third years of the disinflation since the negative output gap has a slightly greater affect on the inflation rate than in the model where the original sacrifice ratio is restored. This results in a smaller trough in output and a decreased cumulative output gap. If left unchanged, with the introduction of the perceived target the cumulative output gap declines by around 0.2 percentage points (from 3.0 to 2.8).

The changes to model dynamics introduced by the inclusion of the perceived target are more evident in a demand shock. This is shown in Figure 19. In both calibrations, inflation does not trough as low, and the monetary response needed to offset the shock is again reduced. For this reason, the cumulative output gap associated with a negative demand shock is greater, since less of it is offset by the monetary authority. This result carries over to other non-policy shocks: the perceived target acts as an anchor to expectations and in general inflation is less affected by shocks to the economy. Output gaps (whether positive or negative) tend to be sustained for longer, however, since the monetary authority does not need to offset as much of the shock.

Unlike the disinflation, in a demand shock the model where the sacrifice ratio associated with a disinflation has been restored shows a greater difference from the base model than the calibration in which no further adjustments are made. This is because the sacrifice ratio is adjusted in the former calibration by reducing the coefficient on the output gap within the price equations. This means that a given shock to demand has less of an impact on inflation.

These calibrations should be regarded as a preliminary attempt at incorporating credibility. Considerably more empirical work is needed to try and identify whether there has been a structural change in the way in which expectations are formed in the 1990s, and to identify changes in key stylized facts used in the calibration of the model, such as the cumulative output loss associated with a disinflation. It is also very important to study the implications of the perceived target using real-world data, rather than in the artificial environment. A good test of the model, will be to simulate the model over history and see how changing the weight on the perceived target alters the extent to which model expectations track other measures of expectations such as those coming from the Conference Board Survey of Forecasters. Another area of research will be to simulate the model in a stochastic environment. In such simulations, factors such as the variance of forecast errors over time can be compared to those from survey data. This will help establish the suitability of different calibrations. It will also be important to conduct sensitivity analyses based on different benchmarks of credibility in order to better define what we would expect to see in terms of factors such as altered trade-offs between unemployment and inflation or output and inflation. These will help to determine how credibility is changing over time.

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[11] This is done by altering the coefficients on the output gap terms within the price equations.

Conclusion

Many aspects of QPM affect the degree to and speed with which price expectations incorporate changes in the inflation target and deviations from the target due to temporary shocks or inconsistent monetary policy. These include the weights on the model consistent and backward components of expectations and the way in which expectations are incorporated into prices. The characteristics chosen in the overall calibration for such things as the sacrifice and benefit ratios in disinflation and inflation shocks, and the degree of monetary response needed to offset shocks or change the target are also crucial determinants of credibility. If credibility has increased in the 1990s, the current calibration can be characterized as incorporating too few credibility effects.

More explicit credibility effects can be incorporated into QPM by introducing a "perceived target" into price expectations based on the lagged perceived target and model consistent inflation in the future. This approach is preferred to one where the target is introduced directly into expectations, since it provides greater flexibility for calibrating different aspects of credibility and reduces problems associated with inconsistency between the announcement effect and both the monetary authority's actions and the intrinsic dynamics of the model. The perceived target can be thought of as agents' assessment of the policy target being used by the monetary authority, which may or may not be equivalent to the announced policy target. The weight on the perceived target within price expectations captures the proportion of agents who base expectations on policy, or the degree to which the average agent believes policy will determine the price outcome.

Unfortunately, calibrating a credibility effect in the model raises more questions than can currently be answered, based on empirical work. The weights in the perceived target were chosen so that when simulated over history, the perceived target does not fully incorporate short-term movements in inflation, but provides a smooth track which acts as a proxy for agents' medium term inflation expectations. Given the lack of evidence currently available, however, a cautious approach was taken to the overall weighting of the perceived target within expectations. It was given a weight of 0.14, which implies a small decline in the cumulative output gap associated with a disinflation. Given the absence of other benchmarks, it is very unclear what should happen to the cumulative output gap/gain associated with a disinflation/inflation. For this reason two calibrations are shown, one in which these cumulative gaps are allowed to adjust fully with the incorporation of the perceived target, and one in which the original values used to calibrate the base model are restored. Above all, this exercise highlights the need for further empirical research on inflation expectations, and the testing of different calibrations with both real world data, and in a stochastic environment.

Appendix 1: Results of horizon-dependent weights on the target

Figure A1.1

Disinflation shock – time dependent coefficients

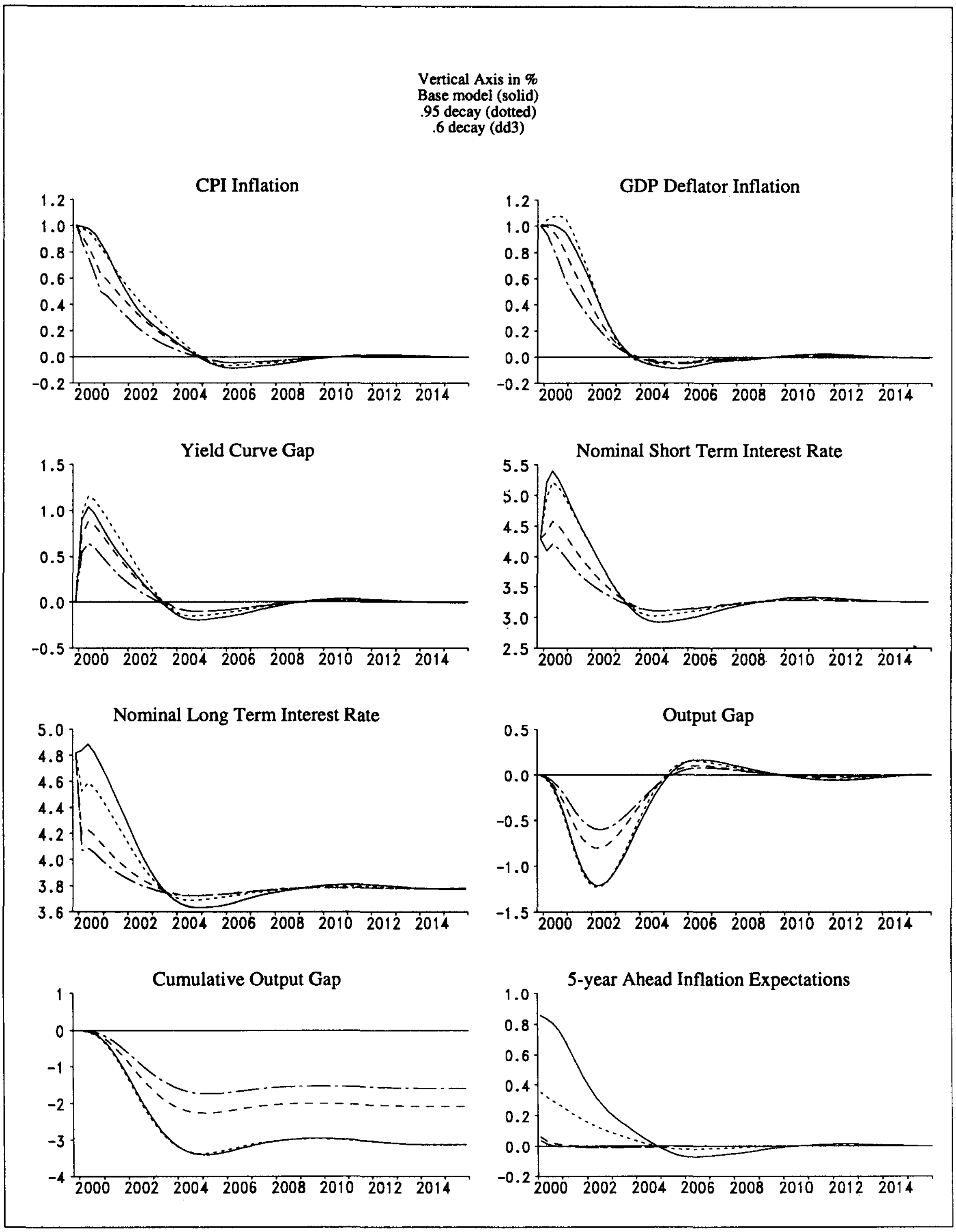
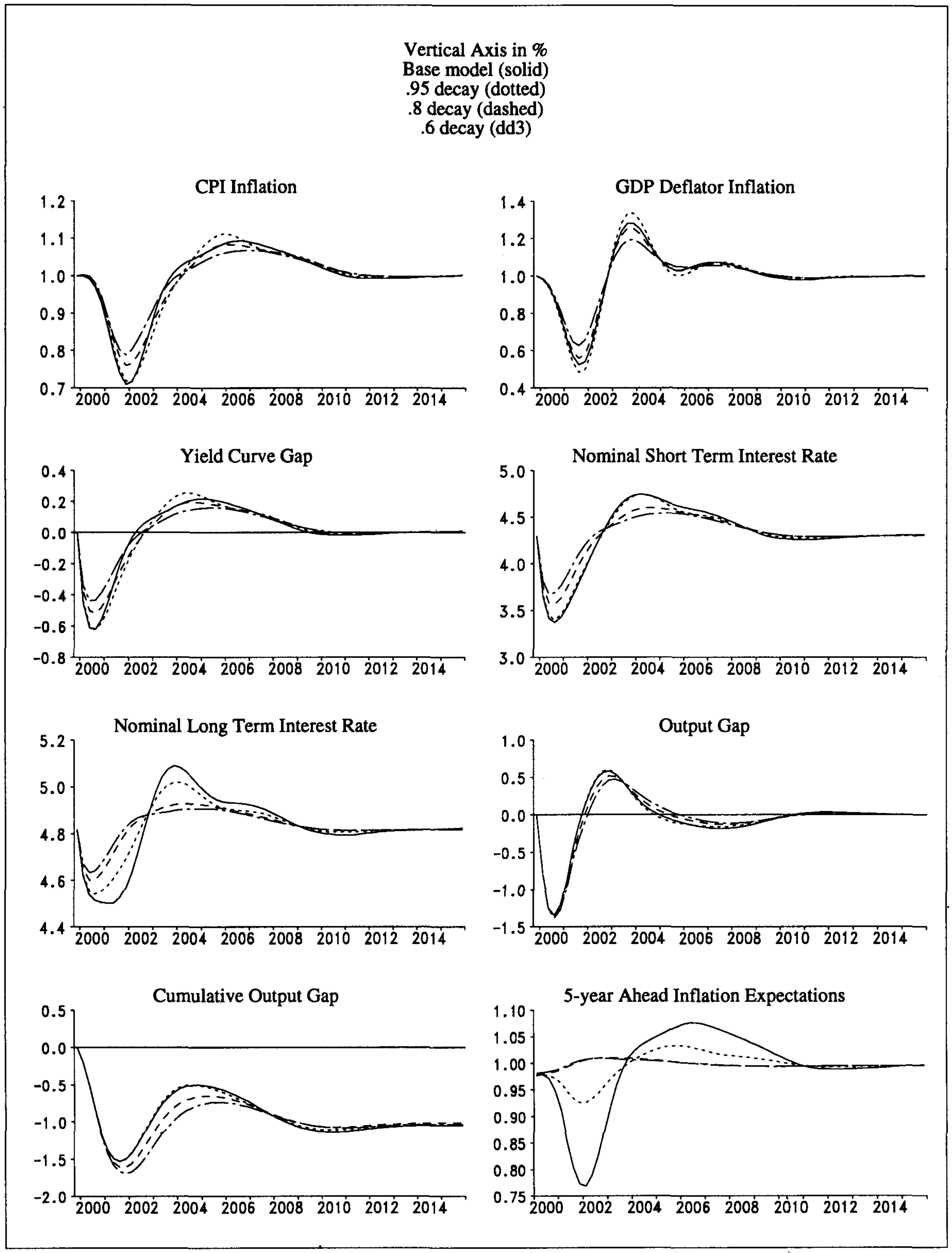


Figure A1.2

Demand shock – time dependent coefficients



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Comments on: "Incorporating credibility in forward looking models: some examples with QPM" by Dinah Maclean

by Gordon J. de Brouwer*

This paper is insightful for a number of reasons. First, it shows that it is important to distinguish between announcement and learning effects in assessing the effect of credibility. Second, the perceived, rather than actual, inflation target matters. Third, it is not necessarily straightforward to deduce credibility effects from the time it takes to get back to the medium-term inflation target after an inflation shock. If there is a high weight on the inflation target in expected inflation, then the gap between actual and target inflation is smaller, and the central bank adopts policy which is less tight than would be the case if there is no credibility. The policy response is smaller but the time it takes to get back to target is similar in both cases. Finally, the paper outlines some of the thinking that led to the way expectations and credibility have been modelled in QPM, showing what worked and what did not, which is always useful to other modellers.

There are a number of other issues which came to mind on reading this paper. The first is that the model is calibrated, and one test of a calibrated model is to see how well it explains history. The calibrated model shows that the perceived inflation target shifted smoothly down in 1988, which is about three years before inflation actually fell in 1991 (Figure 11). According to the model, this had an impact on inflation expectations. But from Figures 15 and 16, inflation expectations did not fall until 1991 when inflation fell, and in fact were trending up slightly in the three years to that time. When the model is used to explain history, it overstates the announcement effect of a change in policy regime. This suggests that we should be sceptical about the credibility effects that come just from announcing regime changes, even when they are carried out.

The effect of credibility is modelled in QPM by changing the behaviour of the forward-looking component of inflation expectations, *not* by increasing the weight on forward-looking expectations. The credibility effect is modelled this way to avoid the greater variability in actual and expected inflation that results when the system is made forward looking and a negative demand shock occurs. But surely it is correct to say that the weight on forward-looking behaviour may also change as a result of credibility. It may be that if the central bank is credible, more people will form expectations in a model-consistent manner. For example, suppose that I do not think that the central bank is serious about its inflation target, such that it will accommodate inflation shocks. Since there is strong persistence in inflation, I expect inflation to be what it has been in the recent past, and I am classed as a "backward looker". If I think the central bank is credible, however, I use what I know about its reaction function in forming my expectations of inflation, and I am now classed as a "forward looker". In fact, I use a model in both cases but the behaviours are classified differently *depending on the credibility of the central bank.* In contrast to the discussion in the paper, this indicates that it is correct to model credibility as having an effect on the classification of behaviour as backward or forward-looking. This still leaves the problem that increasing the weight on the forward-looking component raises short-term inflation variability in QPM. But this is because QPM places the highest weights on very near-term leads of expected inflation. The "problem" can be ameliorated by placing more weight on the medium-term, and less on near-term, leads of expected inflation (de Brouwer and Ellis 1998). Whatever the case, the variability in the system overall falls when expectations are model-consistent.

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The paper goes on to argue that if a central bank has credibility then expectations will change when it announces a new, lower, inflation target. But this need not be the case, for (at least) three reasons:

- the little empirical evidence we have does not support this proposition. For example, the RBNZ's inflation target was changed from 0-2% to 0-3% in December 1996 but expectations of households and key decision-makers did not change when the inflation target was widened. The indexed bond expected inflation series rose but there are measurement problems with these sorts of series and, besides, the series fell back quickly to around 2%. The point is that the change up did not dislodge expectations, so why should a change down?
- credibility may be a function of the stability of a regime, such that changing the regime may be a signal that policy-makers will change the regime again in the future, perhaps in the other direction. Regime changes make other regime changes possible, in the same way that devaluations in a fixed exchange rate regime make further devaluations possible. Moreover, it may be that the change is to a number which is beyond the realm of most people's experience of inflation, and so is regarded as unrealistic and unsustainable, forcing the central bank to run policy even tighter than otherwise (Gagnon 1997). A mean target of 1 to 2% or 2 to 3% may be credible while a mean target of 0 to 1% may not.
- if people do not have model-consistent expectations to start with, then more credibility is not going to reduce the sacrifice ratio if policy-makers want to reduce the inflation target. We simply do not know enough about the way expectations are formed to be confident about the announcement effects of a policy regime change. Like most models, forward and backward-looking behaviour in QPM is calibrated to produce impulse response functions which "make sense". Indeed, while QPM is described as a forward-looking model, the weight on backward-looking expectations is very high. It is important to get behind these representations, to get to the "fundamentals" of how expectations are formed, through, for example, learning mechanisms.

Expectations are treated as homogeneous in QPM whereas in reality they probably tend to differ between markets and sectors in the economy. In Australia at least, it appears that the inflation expectations of financial market economists are less biased than those of households. This raises the interesting issue of the effect of heterogeneous expectations. Using a data-consistent open-economy version of the Ball (1997)/ Svensson model (1997), de Brouwer and Ellis (1998) estimate the inflation and output variability properties of various mixes of forward and backward-looking expectations processes. Since the effects of policy (which is set for one nominal rate) depend on the *ex ante* real interest rate, heterogeneous inflation expectations imply more than one real rate and hence differential impacts of policy on the exchange rate and output. When the foreign exchange market is more forward-looking than price and wage setters, the exchange rate overshoots in response to inflation shocks, and inflation variability is smaller than otherwise. This means that greater credibility (if it is synonymous with more forward-looking behaviour) can have differential impacts, depending on whose expectations are affected.

Finally, while the announcement effects associated with credibility should not be overstated, broader evidence for credibility should not be under-rated. For Australia, there are (at least) three recent pieces of evidence in support of the proposition that the central bank has credibility. The first is that the exchange rate now systematically appreciates when inflation comes in higher than economists in the financial markets expected, which suggests that the foreign exchange market thinks that real interest rates will rise – that is, policy will respond. The second is that wage-setters – be they unions, the institutions which periodically review minimum wages, employers and the government – now expect that interest rates will rise if aggregate wage movements are out of kilter with the inflation target. In other words, there is strong evidence that decision-makers now take the Reserve Bank's inflation target into account when forming their plans. The third is that expectations are broadly consistent with the inflation target, although some caution is required when comparing near-term inflation expectations with a medium-term inflation target.

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Long-run inflation expectations and monetary policy

Antulio N. Bomfim and Flint Brayton*

Introduction

Macroeconomic models are frequently used to simulate the transitional aggregate dynamics that are set into motion by a shift in monetary policy to alter the rate of inflation. A standard result is that the cost of lowering (or raising) the rate of inflation – the integral over time of the deviation of unemployment from its path in the absence of the policy change – varies with how quickly inflation expectations adjust. The more sluggish are expectations, the larger is the unemployment cost per unit of inflation change.

In the Federal Reserve Board's FRB/US macroeconomic model, expectations of long-run inflation play an important role in inflation dynamics. Several different simulation options for the formation of these expectations are available, and as described by Bomfim, Tetlow, von zur Muehlen and Williams (1997), the model's estimate of the unemployment "sacrifice ratio" associated with a change in inflation is affected significantly by the particular expectations mechanism selected. Up until now, however, there has been little empirical basis on which to decide how best to characterize the evolution of long-run inflation expectations. The purpose of this paper is to strengthen the empirical underpinning of this key part of the expectations mechanism in FRB/US by proposing and estimating simple learning rules for the determination of long-run inflation expectations.

Given that inflation in the long run is commonly regarded as a monetary phenomenon, it is natural to look for a connection between long-run inflation expectations and the conduct of monetary policy. Although one might search for evidence of revisions to expectations at times of announcements of policy changes, our prior is that participants in the economy are more likely to scrutinize policy actions more closely than announcements for evidence of a policy shift. Thus, we examine how well various models of learning empirically capture the speed with which long-run inflation expectations respond to a change in monetary policy.[1] The empirical results are then used to construct a version of the FRB/US model in which the expectations held by the private sector about monetary policy are specified as the outcome of learning in a stochastic environment.

A monetary policy regime is typically characterized as a policy reaction function whose structure and coefficients implicitly reflect long-term policy objectives and the speed with which deviations from targets are planned to be eliminated. Changes in policy objectives, including the speed of adjustment, alter the reaction function's coefficients. The models of learning that we examine – rolling regressions and Kalman filtering – yield "real-time" estimates of the coefficients of a posited reaction function. For each approach to learning, the time series of coefficient estimates provides a time series of perceived inflation targets.

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[1] Our linking of long-run inflation expectations to perceptions of the conduct of monetary policy is not the only approach that has been used to characterize long-run inflation expectations. For alternative approaches, see Kozicki, Reifschneider and Tinsley (1996), who present a proxy based on a time-varying intercept in an estimated equation for the rate of inflation, and Kozicki and Tinsley (1996), who describe a measure derived from the term structure of interest rates under the assumption that the real rate of interest is stationary.

The first half of the paper compares available survey data on long-run inflation expectations with our constructed time series of real-time perceptions of the inflation objective of policy. Tentative regression evidence suggests the survey data on long-run inflation expectations is related more closely to our real-time learning constructions than to actual inflation. Nonetheless, the correspondence between the surveys and the constructions is modest, and the most important regressor is the lagged reading from the survey itself. The second half of the paper presents simulations of a version of FRB/US augmented to incorporate several of the real-time learning models. The simulations indicate that some of the learning approaches yield estimates of the unemployment sacrifice ratio that are in accord with the range of conventional estimates.

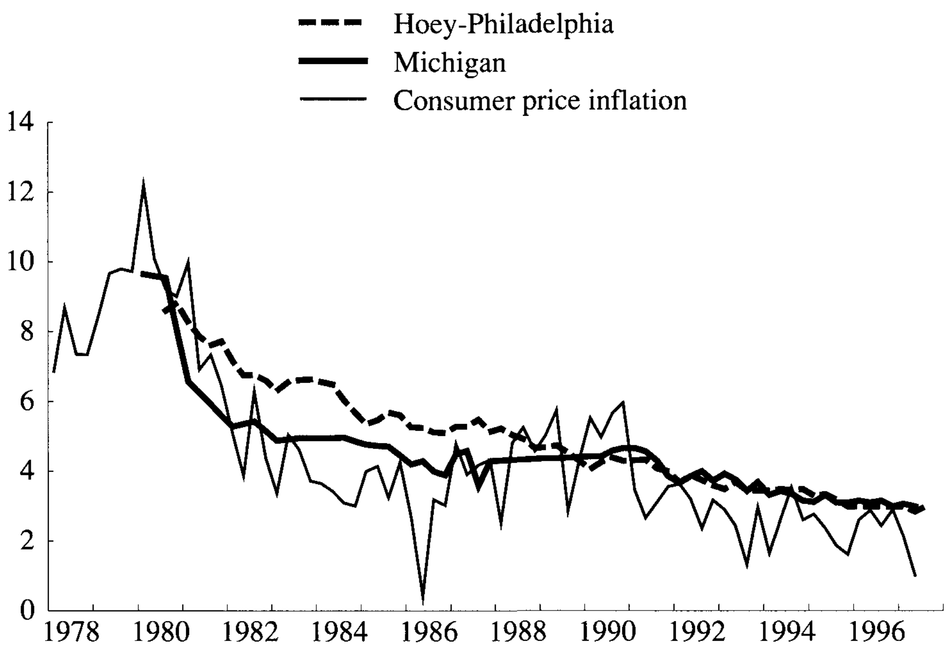
1. Survey measures of long-run inflation expectations

Survey data for the United States on long-run inflation expectations is sparse and available only since 1980. We use two series in our analysis:

- π^s_{mich} is the median inflation expectation over a 5 to 10-year horizon from the Survey Research Center at the University of Michigan. Survey respondents are a random sample of individuals. π^s_{mich} starts in 1980:Q1.[2]
- π^s_{h-p} is a measure of inflation expectations over a 10-year horizon spliced from two surveys. The first segment (1980 through mid-1991) is taken from Richard Hoey's "decision-makers" poll; subsequent observations are from the "Survey of Professional Forecasters" compiled by the Federal Reserve Bank of Philadelphia. π^s_{h-p} starts in 1980:Q3.[3]

Figure 1

Survey data on long-run inflation expectations



[2] Only two observations per year are available from 1980 to 1985 and the series has a gap without observations from 1988:Q1 to 1990:Q1. Missing entries are interpolated linearly. Prior to 1980, a single observation exists for 1979:Q1.

[3] Prior to 1980:Q3, the Hoey survey was also conducted in 1978:Q3 and 1979:Q1.

As shown in Figure 1, π^{s}_{mich} declines fairly rapidly in the early 1980s while the drop in π^{s}_{h-p} is more gradual. The two series converge by 1990 and subsequently edge down in tandem to 3% by 1996. The general consensus holds that monetary policy in the United States shifted in late-1979 to one aiming toward a substantial reduction in the rate of inflation. Neither survey shows a one-time drop in long-run inflation expectations in the immediate aftermath of the policy shift, although, admittedly, the fact that each survey only starts during 1980 makes this conclusion a bit tentative. A question we examine is whether the less-than-immediate response of the two expectations measures is better captured by a learning model in which policy changes become more apparent over time through observation of the changing relationship between the short-run policy instrument and macroeconomic conditions, or whether the survey expectations are simply adjusting to lower inflation as it emerges.

2. A simple model of monetary policy

We assume that historical US monetary policy can be (approximately) represented by an equation for the Federal funds rate in which the explanatory variables are lagged values of the Federal funds rate and current and lagged values of inflation and the deviation of the unemployment rate from an estimate of the natural rate. This specification is closely related to the policy rule proposed by Taylor (1993). While it may be that other macroeconomic or financial factors have influenced policy during certain periods, the posited relationship appears to capture much of the movement in the Federal funds rate since 1966, as long as some variation over time in its coefficients is permitted.

Our starting point is a general dynamic specification in which the Federal funds rate (i) depends on four quarterly lags of the funds rate and the current and first three lagged values of both consumer price inflation (π) and the unemployment gap ($\tilde{u}$):[4]

$$i_t = \alpha + \sum_{i=1}^{4} \beta_i i_{t-i} + \sum_{i=0}^{3} \gamma_i \pi_{t-i} + \sum_{i=0}^{3} \delta_i \tilde{u}_{t-i} \tag{1}$$

Given a set of parameter estimates, the rate of inflation desired by policymakers (π^*) can be calculated as

$$\pi^* = \left(\alpha - \left(1 - \sum \beta\right) r^*\right) / \left(1 - \sum \beta - \sum \gamma\right) \tag{2}$$

if it assumed that (i) the long-run real rate of interest (r^*) is a known constant and (ii) that the equilibrium nominal rate of interest moves one-for-one with equilibrium inflation. Note that the standard inflation stability condition associated with policy rules such as this – that the nominal funds rate change more than one-for-one with changes in inflation – is equivalent to the denominator of equation (2) being negative. If on the other hand, the denominator is zero and interest rates move one-for-one with inflation, monetary policy has no particular inflation target and accepts the current rate of inflation, whatever it is.

Not surprisingly, the coefficients of equation (1) are unstable over time. Figure 2 makes this point graphically. The series plotted are sequences of sub-sample tests for coefficient stability of a reaction function estimated from 1966:Q1 to 1996:Q4. The tests statistics, which are shown for every quarter in this span except the very beginning and end, are reported as ratios to the 5% critical value, and thus numbers greater than 1.0 represent rejections of stability at this significance level. Two test sequences are plotted, one for equation (1) and a second for a version of the reaction function whose dynamic structure has been simplified to eliminate insignificant regressors,

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4 Simultaneity bias is not an issue in estimating this relationship if the reasonable assumption is made that inflation and unemployment are unaffected contemporaneously by the Federal funds rate.

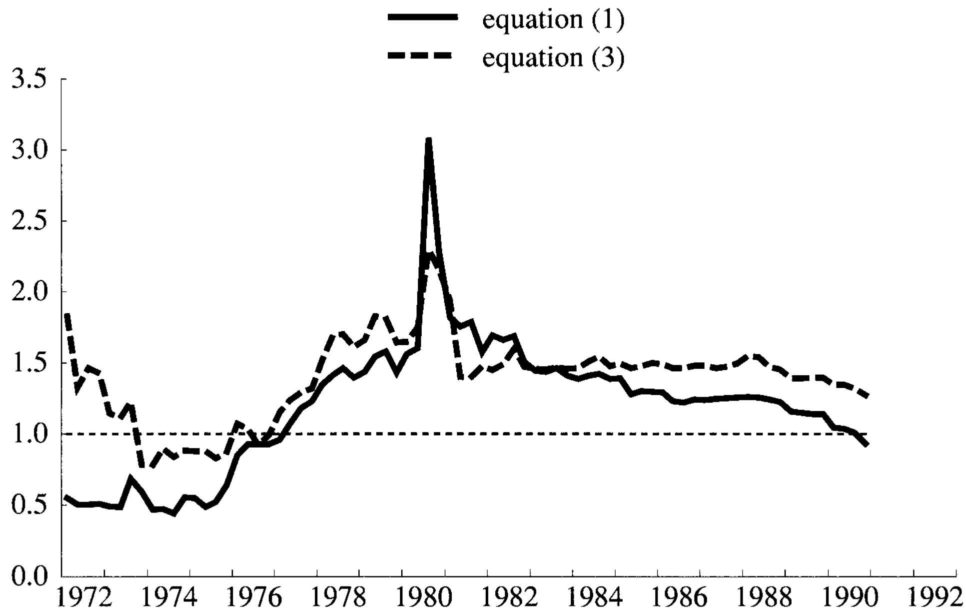
$$i_t = \alpha + \sum_{i=1}^{3} \beta_i i_{t-i} + \gamma \sum_{i=0}^{3} (.25\pi_{t-i}) + \sum_{i=0}^{2} \delta_i \tilde{u}_{t-i} \tag{3}$$

From here on, equation (1) will be referred to as the "long-lags" reaction function and equation (3) as the "short-lags" variant.

Figure 2

Chow test sequences for coefficient stability

Ratio of test statistic to 5% critical value



Although the figure seems to reveal that instability of the reaction function coefficients is pervasive, in fact simply permitting the constant to shift at the end of 1979 leads to a much more stable result.[5] The particular dating of the intercept shift was not chosen on any statistical grounds; rather, the selected switch-point conforms to the commonly held view that monetary policy changed at that time to one aiming to reduce the rate of inflation. Based on the formula given in (2) and the assumption that the real rate of interest is 2%, the "short-lags" specification with an intercept shift indicates that the target rate of inflation fell 4 percentage points from about 6½% in the period up

Table 1

Estimates of the target rate of inflation

Equation	1966-79			1980-96		
	value	95% range		value	95% range	
Long lags	6.44	5.37	7.54	2.75	1.42	3.81
Short lags	6.63	5.45	7.88	2.50	0.99	3.71

5 For the "long-lags" equation, the test statistic for structural change at 1980:Q1 has a p-value of 0.002 when the null hypothesis includes a constant intercept and a p-value of 0.10 when the null includes a shifting intercept. The corresponding p-values for the "short-lags" equation are 0.002 and 0.17.

through 1979 to 2½% since then.[6] The estimates from the "long-lags" version are similar. Confidence ranges around these values are fairly wide and, at the 95% level, encompass values more than 1 percentage point higher or lower than the point estimates.

3. Modeling long-run inflation expectations

We now turn to the question of what someone knowing the general form of the Federal funds rate reaction function could have deduced about policymakers' inflation objectives at different points of time. Hindsight enables the identification of a shift in the inflation target of policy at the end of 1979, but at the time sorting out exactly how policy was changing was undoubtedly difficult. For example, clear identification of the policy change as a lowering of the inflation objective, a more aggressive response to deviations of actual inflation from its target, or some combination of the two, was probably not possible immediately.

Three real-time approaches to estimating the policy reaction function are employed to construct time series of hypothetical perceptions of the policy objective for inflation. The first uses rolling regressions having an estimation interval (window) of fixed length, the second uses rolling regressions in which the estimation interval expands over time but data observations are given less weight as they recede from the end of each estimation period, and the third is the Kalman filter. In each case, the perceived inflation target for any particular quarter is calculated according to equation (2), using the real-time estimates of the reaction function coefficients for that date.

The Kalman filter is the optimal estimation approach when the reaction function coefficients are believed to vary over time as random walks. The first two are more ad hoc in design, though one can think of the optimal window length or decay rate for the rolling regressions as balancing the cost of slower identification of a policy shift as the window lengthens or the decay rate diminishes against the risk of falsely identifying a policy shift when the past is "forgotten" too quickly. The rolling regression approach with declining data importance weights shares one desirable feature with the Kalman filter: Each updates the reaction function coefficient vector in proportion to the gap between the observed value of the Federal funds rate and the value predicted on the basis of the prior estimate of the coefficients. No revision is made to the coefficient vector if there is no surprise to the funds rate.

3.1 Rolling regressions

Rolling regressions were estimated for a variety of window lengths and decay rates. Figure 3 shows the constructed perception of the inflation target derived from the rolling estimation of the "short lags" equation with a 15-year window. This window length yields a constructed series that matches the general pattern of the two inflation surveys somewhat better than do series based on other window lengths. The dotted lines in the figure represent a 95% confidence band around the rolling-regression estimate. Most of the observations from the surveys lie well inside these bands, with the only exceptions occurring at the beginning of the period when some of the high initial survey values lie above the confidence band. Initial values in 1980 of the constructed series, about 7% expected inflation, lie below the survey responses which range between 8 and 10%, and both surveys tend to fall more rapidly than does the constructed measure in the early 1980s. Note the confidence band is not shown for 1996 because it becomes very wide. As the high inflation years of the mid-1970s and

[6] Our results are robust to variation in r^*. As can be seen in equation (2), the effect of r^* on π^* depends on the degree of interest-rate smoothing in the policy rule, which we measure as $\Sigma\beta$. In practice, we find that there is substantial smoothing – the sum of the estimated β_i is close to 1 – and the effect of r^* on π^* is rather small.

very early 1980s gradually fall out of the rolling regression sample, the estimates of the coefficients of the divisor in the formula for the constructed inflation target become less precise.

Figure 3

Perceived inflation objective: rolling regressions with 15-year window

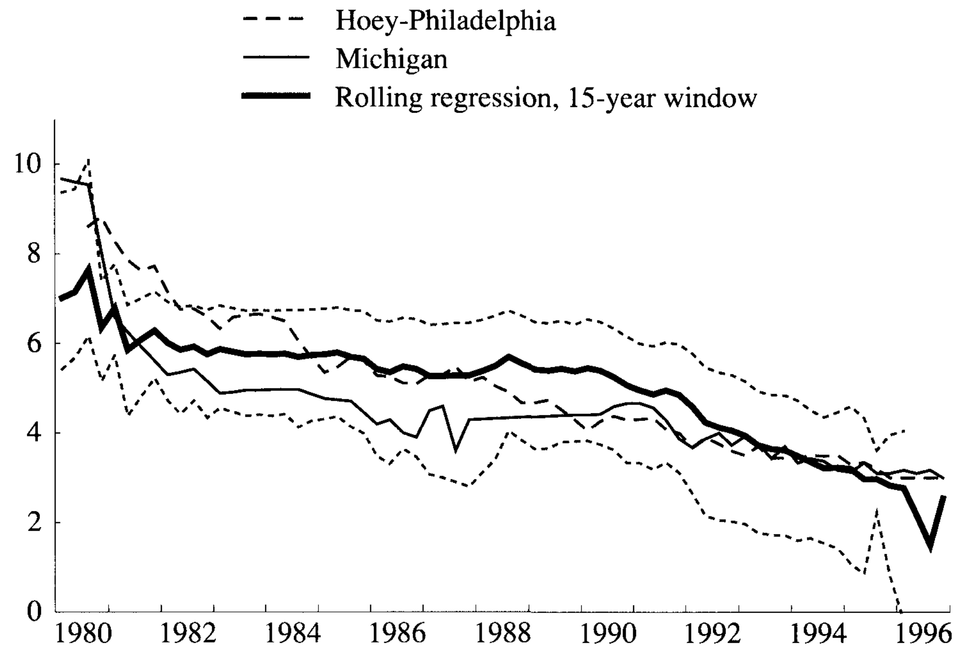
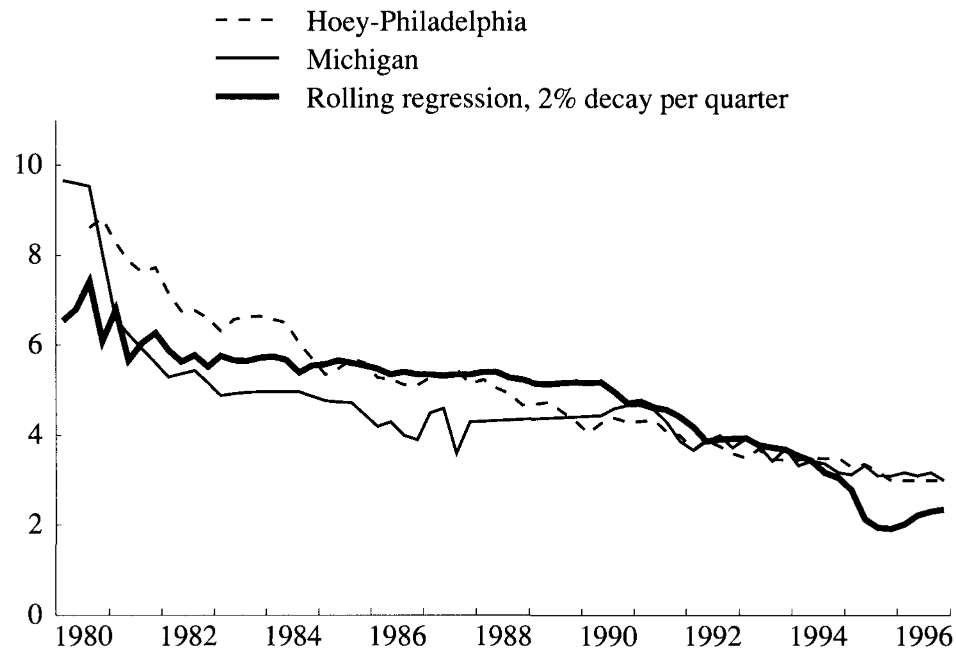


Figure 4

Perceived inflation objective: rolling regressions with 2% decay per quarter



An alternative to the fixed-window regression approach is one in which the estimation sample expands over time but observations are given relatively less weight are they recede into the past. Figure 4 shows a constructed measure of perceived long-run inflation derived under the assumption that the data importance weights decline 2% per quarter. The properties of this series are generally similar to those of the series based on the rolling regression with the 15-year window.

3.2 Kalman filter

To illustrate the Kalman filter approach, we start with a simple Taylor-like policy function

$$i_t = \theta_1 i_{t-1} + \theta_2\left(\bar{\pi}_t - \pi_t^*\right) + \theta_3 \tilde{u}_t + \theta_4 \tilde{u}_{t-1} + \theta_5\left(r^* + \bar{\pi}_t\right) + e_t \tag{4}$$

where $\bar{\pi}_t$ is a 4-quarter moving average of inflation, e_t represents i.i.d. shocks to the reaction function, and $\theta_1 + \theta_5 = 1$.[7]

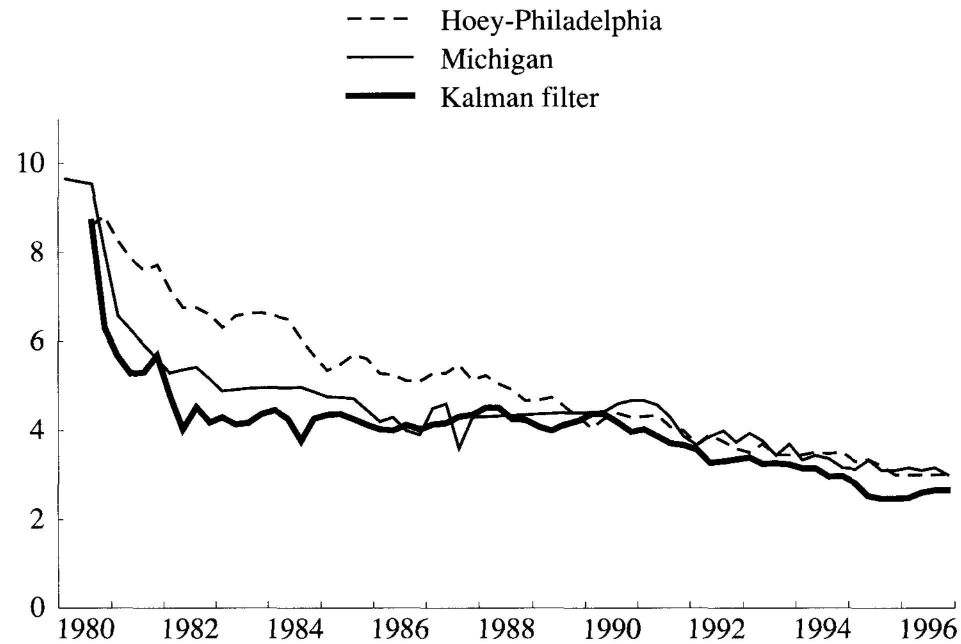
Consider now a framework where the private sector knows the functional form of the reaction function, but not its coefficients or the potentially time-varying inflation target. Agents use a recursive least-squares algorithm to estimate the θ_i parameters and assume that the unobserved inflation target follows a random walk.

$$\pi_t^* = \pi_{t-1}^* + \varepsilon_t \tag{5}$$

where we assume that ε_t is white-noise and uncorrelated with e_t. This specification of the reaction function differs from the one used for the rolling regression approaches in that only one parameter of the policy rule, the inflation target, is explicitly assumed to vary stochastically.

Figure 5

Perceived inflation objective: Kalman filter



The private sector's learning problem can be summarized by the following state-space form:

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[7] It is straightforward to see that equation (4) can be mapped into equation (3).

$$i_t = x_t'\Gamma_t + e_t \tag{6}$$

$$\Gamma_t = \Gamma_{t-1} + \eta_t \tag{7}$$

where $x_t \equiv \left[1, i_{t-1}, \bar{\pi}_t, \tilde{u}_t, \tilde{u}_{t-1}\right]'$, $\Gamma_t \equiv \left[\theta_5 r^* - \theta_2 \pi_t^*, 1-\theta_5, \theta_2+\theta_5, \theta_3, \theta_4\right]'$, and η_t has zeroes everywhere, except for its first entry. Thus, given equations (6) and (7), agents update their estimates of the policy parameters (θ_i) and the inflation target (π_t^*) as each new quarter of data becomes available. The results are summarized in Figure 5. The thick solid line in the figure corresponds to the private sector's perceived inflation target under standard Kalman filter (KF) learning. The constructed series is broadly consistent with the survey data, tracking the Michigan series particularly well. In contrast to the learning mechanisms based on rolling regressions, the generated series drops quite rapidly in the early eighties, from about 9% in 1980 to near 4% in 1982 – this decline is comparable to the one registered by the Michigan survey, but faster than suggested by the Hoey-Philadelphia data. Turning to more recent readings, the standard KF-based learning algorithm places long-run inflation expectations at about 2½% in early 1997, about 50 basis points below both surveys.[8]

3.3 Are the models consistent with the surveys?

Simple regressions are used to characterize more formally the relationship between the inflation surveys and the constructed series,

$$\pi_t^s = \alpha_0 + \alpha_1 \pi_{t-1}^s + \alpha_2 \pi_t^c + \alpha_3 \pi_t, \tag{8}$$

in which one of the inflation surveys (π^s) is regressed on a constant, its own lag, a constructed target inflation perception (π^c), and actual inflation (π).

Table 2 reports a pair of regressions for each combination of the two inflation surveys and the three constructed inflation perceptions presented above. For each combination, the first regression restricts the intercept to zero and the sum of the other coefficients to be one, while the second regression is unrestricted. For the Michigan inflation survey (π_{mich}^s), coefficients on the perceived inflation target are uniformly larger and statistically more significant that are coefficients on actual inflation.[9] Indeed, the coefficients on the perceived targets tend to be highly significant while most coefficients on actual inflation are insignificantly different from zero. Nonetheless, the most important regressor is the lagged value of the survey, whose coefficient ranges between 0.7 and 0.8.

Qualitative aspects of the regressions for the Hoey-Philadelphia survey (π_{h-p}^s) are similar to those for the Michigan survey. The perceived inflation targets tend to be more significant than is actual inflation and the survey data are quite inertial. Quantitatively, for the Hoey-Philadelphia survey, the degree of difference in significance of the perceived target and actual inflation is reduced, and coefficients on the lagged survey observation are even higher.[10]

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[8] We have experimented with a version of the Kalman filter approach that allows for subsample variation in r^* and found that our results are little changed.

[9] Estimation results in Table 2 are little affected if the inflation perceptions are entered with a lag rather than contemporaneously, if the first lag of actual inflation is used in place of its current value, or if lags 0 to 3 or lags 1 to 4 of actual inflation are entered. Furthermore, additional regressions indicate that Granger causality runs from the perceived targets to the Michigan survey but not vice versa.

[10] Another difference is that the Hoey-Philadelphia survey and the perceived inflation targets each Granger causes the other.

Table 2

Regressions of survey inflation on constructed series

Survey inflation	Constructed inflation	Coefficient: cnst	π^s_{t-1}	π^c_t	π_t	Regression standard error
π^s_{mich}	w = 15	–	0.80	0.14	0.06	0.304
			(14.2)	(3.5)	(1.8)	
		0.39	0.72	0.14	0.02	0.263
		(2.7)	(13.9)	(3.3)	(0.7)	
	d = 0.02	–	0.79	0.16	0.05	0.303
			(14.5)	(3.9)	(1.6)	
		0.44	0.73	0.13	0.02	0.265
		(3.2)	(14.0)	(3.2)	(0.8)	
	Kalman	–	0.71	0.34	-0.04	0.247
			(17.2)	(8.0)	(1.5)	
		0.28	0.58	0.39	-0.00	0.238
		(2.1)	(9.7)	(5.2)	(0.1)	
π^s_{h-p}	w = 15	–	0.89	0.07	0.04	0.184
			(30.7)	(2.3)	(2.9)	
		0.09	0.91	0.05	0.01	0.180
		(0.9)	(30.5)	(1.3)	(0.5)	
	d = 0.02	–	0.87	0.10	0.053	0.178
			(32.2)	(3.3)	(2.2)	
		0.08	0.89	0.07	0.01	0.176
		(0.9)	(29.8)	(2.0)	(0.4)	
	Kalman	–	0.91	0.09	0.00	0.175
			(58.7)	(3.8)	(0.1)	
		-0.03	0.90	0.11	-0.00	0.178
		(0.2)	(29.8)	(1.7)	(0.1)	

Notes: t-statistics are shown in parentheses. Michigan (π^s_{mich}): sample period is 1980:Q3 – 1996:Q4. Hoey-Philadelphia (π^s_{h-p}): sample period is 1981:Q1 – 1996:Q4.

On the whole, the regression tests support the view that long-run inflation expectations move, in part at least, with changing perceptions about monetary policy objectives as derived from learning models. The adaptive expectations view has less support.

4. Simulation analysis

The historical analysis suggests that for FRB/US simulations which assume that the private sector has incomplete information about the nature of monetary policy, gradual adjustment of long-run expectations might be better specified as the outcome of learning about the parameters of the policy reaction function than as a simple partial adjustment to observed inflation. The remainder of this paper describes some exploratory FRB/US simulations in which the evolution of long-run inflation expectations is modeled as the outcome of learning. Because learning is a process of signal extraction, a realistic analysis of alternative learning mechanisms requires a stochastic rather than deterministic simulation environment.

4.1 Long-run inflation expectations in FRB/US

As mentioned in the introduction, FRB/US has several different simulation options for expectations formation. The option employed here bases expectations on the forecasts of a VAR system that at its core has a set of three equations for the Federal funds rate, inflation and the output gap. The VAR system is restricted so that as the planning horizon lengthens, period-by-period inflation expectations approach the long-run inflation expectation.[11] The long-run inflation expectation is an anchor or "endpoint" that at any point of time is predetermined in the calculation of expectations having a shorter horizon. In FRB/US simulations, the manner in which the inflation expectations endpoint moves over time has up until now been specified as either adaptive, in the sense of adjusting gradually toward actual inflation, or as embodying full knowledge of the true long-run policy objective for inflation. The simulations reported next instead use one of the regression-based learning algorithms.

4.2 Design of stochastic simulations

The FRB/US model consists of about 40 stochastic equations, numerous identities and about 100 exogenous variables. For stochastic simulations, equations are added for 10 key exogenous variables, such as the price of oil, so that they can be easily given random shocks. For the 50 stochastic equations in the augmented model, shocks are bootstrapped from historical residuals. In each period simulated, a historical quarter between 1966:Q1 and 1995:Q4 is randomly chosen and the vector of equation residuals associated with that quarter is drawn. Most FRB/US equations have residuals that are serially uncorrelated. For a few financial equations, however, residuals are serially correlated, and AR(1) error-propagation equations are added to the model in these instances. Monetary policy is characterized by the version of the "short-lags" Federal funds rate reaction function estimated from 1966:Q1 to 1996:Q4 that allows for a shift in its intercept at the end of 1979.

Several special issues arise in stochastic simulations that incorporate learning algorithms such as those discussed above. One is the need for initial conditions from which to start the algorithms. Although the last 15 or 20 years of US macroeconomic data could serve this purpose, with the stochastic simulations running from the present out into the future, other considerations make it easier to start from a deterministic baseline characterized by steady-state balanced growth. For this reason, the stochastic simulations reported in this paper have an initial 15-year period in which the long-run inflation expectation is exogenous. Then, with a long enough simulated "history" available, one of the learning algorithms is switched on for the remaining 35 years of each simulation.

A second issue concerns the shocks applied in the stochastic simulations to the Federal funds rate reaction function. The historical residuals of this equation are quite variable – the standard deviation is about 100 basis points – and include several outliers. An important question is to what degree these residuals, especially the outliers, represent actual surprises to participants in the economy and to what degree they reflect well-understood responses of the funds rate to special short-run factors that are not included in the reaction function. One example is the credit-control episode of 1980, which led to large short-run gyrations in GDP and interest rates as well as large residuals to the estimated equation for the Federal funds rate. Outside of a few episodes, however, it is more difficult to gauge the appropriate magnitude of "true" errors. In the stochastic simulations, two modifications are made to the funds rate shocks. First, to reduce the influence of outliers, the residuals are drawn from a normal distribution rather than from the historical set residuals. Second, to examine the sensitivity of simulation results to the magnitude of funds rate shocks, the standard deviation of the normally-distributed shocks is chosen alternatively as one-half or the same as the standard deviation of the historical residuals.

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[11] The role of long-run expectations in the FRB/US model is discussed in more detail in Brayton, Mauskopf, Reifschneider, Tinsley and Williams (1997) and Bomfim and Rudebusch (1997).

Finally, a metric that will be used to evaluate the performance of the alternative learning procedures is the unemployment sacrifice ratio associated with a monetary policy shift that aims to reduce the rate of inflation one percentage point. For each particular learning procedure analyzed, two sets of stochastic simulations are run, one set in which the policy target for inflation is constant over time and a second set in which a one percentage point reduction in the inflation target is introduced in year 20. Because the same sequence of shocks is drawn in each stochastic set, pairwise comparisons of individual simulations can be made. Each simulation set consists of 50 replications.[12]

4.3 Simulation results

Table 3 summarizes the results of the FRB/US stochastic simulation experiments. Each row corresponds to a particular experiment whose design is described in the left pair of columns. The middle four columns report the standard deviations of key macroeconomic variables from the set of stochastic simulations in which the policy inflation target is held constant, and the three columns on the right present statistics based on comparing the disinflationary set of stochastic simulations with the set having the constant inflation target.

Table 3

Stochastic simulation results

Simulation design [1]		Statistics under constant π^{*} [2] (standard deviation)				Disinflation statistics [3]		
						Sacrifice ratio		Years for π^e
π^e	ε_i	π	π^e	$\tilde{x}$	i	median	s.e.	to fall 0.9
d=0.04	1.0	1.85	1.12	3.01	2.88	2.88	1.20	11.50
d=0.03	1.0	1.79	0.85	2.98	2.79	2.92	0.95	12.50
d=0.02	1.0	1.74	0.56	2.94	2.70	3.52	0.77	12.75
d=0.05	0.5	1.71	0.72	2.83	2.45	2.51	0.75	8.50
d=0.04	0.5	1.67	0.52	2.79	2.40	2.69	0.80	9.75
d=0.03	0.5	1.65	0.37	2.78	2.35	3.02	0.60	12.50
d=0.02	0.5	1.64	0.24	2.77	2.31	3.69	0.59	19.75
w=10	1.0	1.88	1.18	3.08	2.99	2.97	2.00	8.25
w=15	1.0	1.74	0.64	2.93	2.71	3.39	0.79	14.75
w=8	0.5	1.74	0.86	2.85	2.55	2.64	0.98	8.50
w=10	0.5	1.68	0.55	2.81	2.39	2.64	0.97	10.75
w=15	0.5	1.64	0.30	2.76	2.30	3.43	0.73	13.00
kf	1.0	1.79	0.70	2.96	2.79	3.09	1.03	12.00

[1] Rolling-regression learning is denoted by w (= years) for fixed window or d (= decay rate) for expanding sample with declining data weights. "*kf*" denotes the Kalman filter. ε_i is the scale factor applied to shocks to the Federal funds rate.

[2] Standard deviations for quarterly observations on inflation (π), the long-run expected inflation(π^e), deviation of output from potential ($\tilde{x}$) and the Federal funds rate (i). Standard deviations are calculated from years 20-50.

[3] Sacrifice ratio statistics are from year 30 of the simulations.

[12] Occasionally, the regression-based learning algorithms calculate a value for the long-run inflation expectation that is wildly high or low, probably when the learning algorithm estimates the reaction function coefficients imprecisely. To prevent rare occurrences of this sort from having a large impact on a particular simulation, upper and lower boundaries are placed on the permissible values of the long-run inflation expectation. The upper limit is 3 percentage points higher than the policy target for inflation in the constant inflation simulations and the lower limit is 3 percentage points lower than the policy target in the disinflation simulations.

Focusing first on the rows that correspond to the rolling regression approaches which performed best in the historical analysis (d=0.02 and w=15), the median unemployment sacrifice ratio is close to 3½, a value which is well above current "consensus" estimate of 2 or so – see, e.g. Ball (1994). This finding is not affected if the magnitude of the shocks to the funds rate equation is scaled down by one-half. These particular parameterizations of the rolling regressions also result in disinflations that are very slow: The average length of time it takes for the long-run inflation expectation to fall 0.90 percentage points, or 90% of its ultimate decline, is well over 10 years. Because the process of learning about the change in the policy objective for inflation occurs very gradually, and the degree of inertia in expectations is high, a sizeable increase in the unemployment rate is required to lower the rate of inflation.

The learning process is accelerated and the median sacrifice ratio reduced by adjusting the rolling regression parameters to speed the rate of decline of the data importance weights or shorten the estimation window. When the shocks to the Federal funds rate have their full historical variability, raising the decay rate to 4% per quarter or shortening the regression window to 10 years reduces the simulated sacrifice ratio to 3. If the funds rate shocks are reduced to one-half their historical variability, a decay rate of 5% or a regression window of 8 years results in a sacrifice ratio of a bit more than 2½.

Up to this point, the shortening of the actual or effective length of the rolling regressions has little deleterious effect on measures of macroeconomic volatility when the policy holds the target rate of inflation fixed. Increases in the standard deviations of inflation, output and the Federal funds rate are minor. Note, however, that the variability of the sacrifice ratio across individual simulations gets substantially larger in some cases. Further contraction of length of the rolling regressions leads to much higher volatility of expected long-run inflation and this spills over into higher macroeconomic variability.

The last two rows of Table 3 summarize the results of stochastic simulations based on Kalman filter learning. As shown in the middle columns, the measures of macroeconomic volatility reported in the table are little affected by allowing the FRB/US agents to use the Kalman filter to form long-run inflation expectations. The same is not true for the disinflation statistics: The median value of the sacrifice ratio is about 3, roughly ½ point lower than the sacrifice ratios implied by the rolling regressions that performed best in the historical analysis, but still on the high end of estimates reported in the empirical literature.

Concluding remarks

Long-run inflation expectations play an important role in the short-run macroeconomic dynamics of the FRB/US model. Yet, FRB/US has so far lacked an empirically based approach to modeling the evolution of long-run inflation expectations when the private sector is uncertain about the ultimate inflation goals of the policymaker. In its narrowest sense, we view this paper as an attempt to fill in this gap. The learning schemes estimated here come from estimated learning processes that are based on explicit beliefs about monetary policy and macroeconomic conditions. In a broader sense, this paper illustrates how real-time learning can be incorporated into a large-scale macroeconomic model in a way that attempts to be both data- and model-consistent.[13] We see this as an important contribution given the two most common modeling alternatives entail assuming either

[13] Other researchers who looked into similar issues include Hall and Garratt (1995), who examined similar issues in the context of the London Business School Model, and Fuhrer and Hooker (1993), who analyzed the economic implications of alternative learning schemes in a small-scale macro model.

that the agents have full and complete knowledge of the workings of the economy – the rational expectations hypothesis – or that agents are limited to passively responding to actual developments in lagged inflation.

We should also emphasize that the nature of our findings extends beyond the interest of large-scale macro modelers. In particular, our method allowed us to estimate survey-independent measures of market participants' long-run inflation expectations. This is of value in and of itself given that the available survey data often cover only a short span of time. More important, armed with our constructed time series, we plan to examine how different learning models conform to historical and perceived conditions in financial markets – e.g., ex-ante long-term real interest rates – and how well they anticipate future developments in inflation.

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Comments on "Long-run inflation expectations and monetary policy" by A. Bomfim and F. Brayton

by Franck Sédillot

This is a very interesting paper because it provides an elegant method to extract real time perceived inflation targets. The paper can be divided into two parts. The first part proposes and estimates simple learning rules for the determination of long-run inflation expectations. The second, following a previous paper presented at the BIS last year,[1] provides an empirical test (applied to the sacrifice ratio) of these learning rules by using the FRB/US model. The remainder of the discussion is devoted mainly to the first point, i.e. the formation and the role of long-run expectations.

What is a learning rule?

Learning can be modelled on the basis of a number of assumptions about the underlying knowledge which agents possess. The assumptions made here are that the agents use a "reasonable" learning rule which remains constant over time to form expectations. In fact, it is assumed that agents know the reduced form of the whole system but do not know some or all of the parameters. The reduced form of this model is a combination of stable structural equations and of time varying parameters of the expectation rules. In this paper the authors assume that US monetary policy can be approximated by a standard monetary policy reaction function.[2] So the short-term interest rate depends on its own past values, consumer price inflation and the unemployment gap (a proxy for the output gap). This is the "common" knowledge of the private agents summed up in equation 1.

$$i_t = \sum_{i=1}^{k} \beta_i i_{t-i} + \sum_{i=0}^{k} \gamma_i \pi_{t-i} + \sum_{i=0}^{k} \delta_i \overline{u_{t-i}} \tag{1}$$

How to extract long-run inflation expectations?

Given the assumption that in the long run the unemployment gap is zero, that the real short-term interest rate is constant (the inflation target and the real interest rate target cannot be separately identified) and that the equilibrium interest rate moves one to one with inflation we can extract, after the estimation of the equation, the target values of the inflation rate (see equation 2).

$$\pi^* = (\alpha - (1 - \Sigma\beta)r^*)/(1 - \Sigma\beta - \Sigma\gamma) \tag{2}$$

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[1] Bomfin, A., R. Telow, P. von zur Muehlen and J. Williams (1997): "Expectation, Learning and the Costs of Disinflation: Experiment using the FRB/US Model". *FEDS working papers*, No. 97-42.

[2] Clarida, R., J. Gali and M. Gertler (1997): "Monetary Policy Rules in Practice: Some International Evidence". *CEPR Discussion Papers*, No. 1750.

So for each quarter from 1980 to 1996, it is possible to build a time series of perceived inflation objectives. Two methods are used: rolling regressions or recursive regressions (in which first data observations are given less weight as they go back from the end of the period) and the Kalman filter. In the first case, the coefficients of equation 1 are allowed to vary over time. In the second case, the coefficients also vary over time but their estimation takes into account the unobservable component exhibited by the Kalman filter. For example, this kind of method has been recently used by Gordon[3] for estimating his time-varying NAIRU. Let's turn now to the results.

What are the results?

Surveys on expected inflation provide a useful "benchmark" for testing the reasonableness of estimated long-run inflation expectations. For the United States, data are almost exclusively concerned with short-term expectations. However, there are two exceptions: a survey of market participants conducted by R. Hoey coupled with a quarterly survey of professional forecasters conducted since the first quarter of 1991 by the Federal Reserve Bank of Philadelphia (these two surveys ask for the average expected inflation over the next ten years) and a survey conducted by the University of Michigan which asks for inflation over the next 5-10 years. The results are quite convincing. Broadly speaking, the properties of the series extracted are similar to those of the surveys. But the extracted perceived inflation objective using the Kalman filter method is almost always under the Michigan survey. In fact, long-run inflation seems to be very close to inflation itself. I will come back to this point further in the discussion. Another way to validate the estimated series is to regress the inflation survey on its own lag, a constant, the constructed inflation perceptions and actual inflation. In all the regressions most coefficients on actual inflation are insignificantly different from zero. I turn now briefly to the simulation results. They show that because of the high degree of inertia in expectations the sacrifice ratio is well above common estimates for the United States (2%). I will also come back to this point further in the discussion.

Some comments

My first comment is a general question about the link between moving endpoints and learning rules. In the FRB/US model, a core VAR is used which represents the aggregate information that is available to all agents in the economy. This core VAR consists of three equations: one for the short-term interest rate, for inflation and for the output gap. The main point is that the long-run level of these three variables is constrained by transversality conditions[4] (broadly speaking a kind of a sophisticated long-run level of the variable). For the short-term interest rate a forward rate is used. From the interest rate endpoint one can then deduce an inflation endpoint (e.g. expected inflation in ten years) thanks to the Fisher relation (see equation 3 and Figure 1 which plots the inflation rate and the inflation endpoint in the case of France).

$$r_t^{\infty} = \rho_t^{\infty} + \pi_t^{\infty} \tag{3}$$

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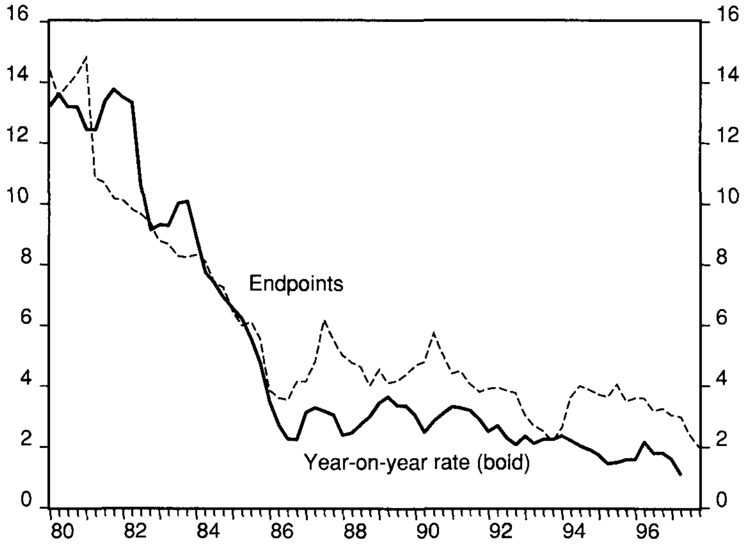
[3] Gordon, R. J (1997): "The Time Varying NAIRU and its Implications for Economic Policy". *Journal of Economic Perpectives,* 11 (winter), pp. 11-32.

[4] Kozicki, S., and P. A. Tinsley (1996): "Moving Endpoints and the Internal Consistency of Agents' ex Ante Forecasts". *FEDS working papers*, No. 96-59.

Figure 1

Inflation and endpoints for inflation in France



In fact, there are two alternatives for constructing an historical measure of the inflation endpoints: one can be based on the interest rate endpoints (which also tracks surveys very well); the other can be based on the learning process. Moreover, the learning rule is not unique. For example Kozicki, Reifschneinder and Tinsley[5] use a learning model which allows for shifts in expected inflation to extract endpoints. They found that the estimated inflation is very close to available surveys. That leads me to my first question. What is the best alternative to extract inflation expectations: the Fisherian approach or the learning rule? What is the difference between your approach and the approach from Kozicki, Reifschneider and Tinsley?

My second question consists in fact of two questions and is closely related to the Kalman filter results. It seems that expected inflation extracted with the Kalman filter technique fits a moving average of consumer price inflation very closely (see the line in Figure 1 above and the bold line in Figure 5 in the paper). If this is the case, is there any reason behind this? Kozicki, Reifschneinder and Tinsley found with their own learning rule that during the period of disinflation, expected inflation overestimates the surveys and so does consumer price inflation. Kozicki and Tinsley also found a high degree of inertia of expectations in the case of a moving endpoints formulation. In the case of France, we also found a similar result (see Figures 1 and 2).

There is a different way to ask this question. In the simulation results the authors found that the disinflation process in the case of the Kalman filter is faster than in the case of the rolling regressions. I think this is because in the case of rolling regressions the expected long run inflation is well above inflation. So to diminish the sacrifice ratio towards 2% the expected inflation has to be close to inflation itself (that is the Kalman's filter result). From my point of view there are two ways to deal with such a problem. Either one considers that the inflation expectations have a high degree of inertia which is consistent with the authors' model and the surveys, but gives disappointing results in terms of sacrifice ratio. Or one considers the inflation expectations close to measured inflation which is less consistent with the model, but gives more convincing results. How is this kind of dilemma solved?

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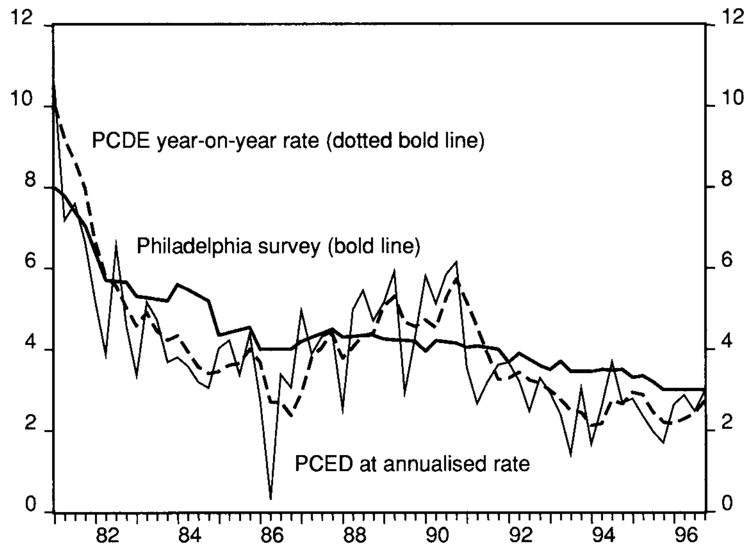
5 Kozicki, S., D. Reifschneider and P. A. Tinsley (1995): "The Behavior of Long-Term Interest Rates in the FRB/US Model". *Mimeo.*

Figure 2

Survey and inflation in the US

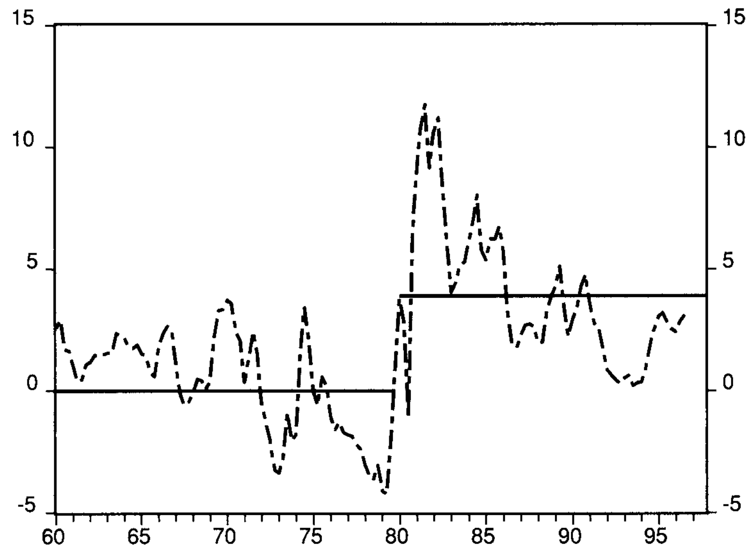


My last question is more technical and relates to the monetary policy reaction function. In the calculation of the inflation target, the authors assume that the real rate is constant and equal to 2%. I think they have proxied the real interest rate target by the sample average of the real interest rate and thus their estimate of π_t^* cannot differ too much from the sample average of π. As we can see in Figure 3, there is a break in the mean of the real Fed funds rate. They allow for this break in their long lag and short lag equations. This stabilises the coefficients of equation 1 and the shift in the intercept raises the level of the real interest rate at the beginning of the 1980s (its mean jumps from 0% to about 4%). But in the Kalman filter equation (see equation 4) they have no intercept. So what level of real interest rate is taken or, in other words, how is the shift in the variable taken into account?

$$i_t = \theta_1 i_{t-1} + \theta_2(\overline{\pi_t} - \pi_t^*) + \theta_3 \overline{u_t} + \theta_4 \overline{u_{t-1}} + \theta_5(r^* + \overline{\pi_t}) + \varepsilon_t \tag{4}$$

Figure 3

Fed funds real rate



Forward interest rates and inflation expectations: the role of regime shift premia and monetary policy

Hans Dillén and Elisabeth Hopkins*

Introduction

In recent years attention has been paid to the use of forward interest rates as monetary policy indicators of inflation expectations, see, for example, Dillén (1996), Svensson (1994), and Söderlind (1995). However, the Swedish experience of using forward interest rates as monetary policy indicators suggest that these rates are troublesome to interpret. The problem is essentially that fluctuations in forward rates often have been highly volatile and that it is difficult to relate these fluctuations to economic factors that usually are believed to affect interest rates, e.g. inflation expectations.

In this paper we argue that there are two major explanations for the relatively high and volatile development of forward interest rates: (i) investors' fears that the economy will switch to a high inflation regime give rise to a fluctuating regime shift premium; and (ii) expectations of monetary policy actions amplify the effect on forward interest rates originating from fluctuations in inflation expectations. In addition to these explanations we also include a time-varying term premium in the analysis. We show in an empirical analysis based on Swedish data the significance of adjusting for regime shift premia and taking the interaction between inflation expectations and expectations of monetary policy actions into account. Term premia are normally small, but occasionally they are of importance. In other words, it is essential to take regime shift premia (and sometimes term premia) and expectations of monetary policy actions into account when forward rates are used as monetary policy indicators.

In the literature it has been recognised that the usefulness of forward interest rates as indicators of future inflation expectations depends on the relative volatility and the correlation of inflation expectations and expected real interest rate. Several studies have tried to extract information about inflation expectations from the term structure. Fama (1990) finds that bond prices contain information about future values of a range of economic variables, such as future spot rates, inflation, real returns and expected term premiums. Mishkin (1990) analyses the information content of the term structure for future inflation and finds that nominal interest rates with maturities of nine to twelve months contain information about future inflation. In Söderlind (1995) "the forward rate rule", which states that all movements in forward interest rates reflect fluctuations in expected inflation, is evaluated on US and UK data. He finds that this rule performs reasonably well. Regime shift effects in the term structure were originally analysed by Hamilton (1988) but in a somewhat different context. This study is instead related to the kind of inflation regimes analysed by Dillén (1997), Evans and Wachtel (1993) and Evans and Lewis (1995).

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The outline of the paper is as follows: Section 1 is an overview of issues and concepts we aim to address in this study. In Section 2 the regime shift model is presented as well as a characterisation of the regime shift premium. Section 3 presents the estimates of regime shift premia in Sweden, which is an extended update of the estimates provided by Dillén (1996). The relation between inflation expectations obtained from surveys and forward interest rates is discussed and estimated in Section 4, which also includes an analysis of the role of monetary policy expectations in this context. The final section concludes.

1. Issues and concepts

The purpose of this study is to develop methods useful to extract and interpret information contained in forward interest rates, especially inflation expectations and expectations of future monetary policy. In this respect it is interesting to analyse *forward excess returns*, η, defined as the difference between forward interest rates and the future short-term interest rate:

$$\eta(t,\tau) = f(t,\tau) - i(t+\tau) \tag{1}$$

where $f(t,\tau)$ is the forward interest rate[1] at time t with settlement at time $t+\tau$, and $i(t+\tau)$ is the short-term interest rate at time $t+\tau$. Under the pure expectations hypothesis the excess forward return can be seen as a forecast error with $E_t[\eta(t,\tau)] = 0$, i.e., the forward interest rate is an unbiased indicator of the future short-term interest rate. Moreover, under the Fisher hypothesis the future short-term interest rate is the sum of the expected future short-term real interest rate and the expected future inflation rate.

A problem with using forward interest rates as monetary policy indicators is that they are probably affected by factors which make it difficult to extract and interpret information of expectations of future economic conditions. Forward interest rates may contain a time-varying term premia, $\rho(t,\tau)$, (such that $E_t[\eta(t,\tau)] = \rho(t,\tau)$) which makes it more troublesome to extract information about expectations. Moreover, it is often desirable to separate different types of expectations. There might be Peso type problems, i.e. expectations of drastic, but unlikely, events (for example, devaluations or default) that might have impact on interest rates. Drastic events of this kind are sometimes modelled as regime shifts and in this study we consider the possibility that expectations of a shift to an inflationary regime will cause peso type problems. In other words we will assume that there is a regime shift component in forward interest rates. The above discussion leads us to consider the following decomposition of the forward interest rate:

$$f(t,\tau) = r^e(\tau) + \pi_0^e(\tau) + f_{rs}(t,\tau) + \rho(t,\tau) \tag{2}$$

where $r^e(\tau)$ and π_0^e are the expected short-term real interest rate and inflation rate *within the regime* respectively, $f_{rs}(t,\tau)$ is the regime shift component of the forward interest rate (or the regime shift premium) and $\rho(t,\tau)$ is a time-varying term premium. By construction we have that $E_t[i(t+\tau) \mid \text{no regime shifts}] = r^e(\tau) + \pi_0^e(\tau)$, and regime shift expectations (or peso type problems in general) are represented by the regime shift component, $f_{rs}(t,\tau)$. Thus, the regime shift component will have a systematic effect on excess forward returns as long as no regime shift occurs, i.e.

$$\eta(t,\tau) = f_{rs}(t,\tau) + \rho(t,\tau) + e(t,\tau), \qquad \text{with } E_t[e(t,\tau) \mid \text{no regime shifts}] = 0. \tag{3}$$

1 The interest rate agreed at time t for a short-term loan at time $t+\tau$.

2. Regime shift and term premia

Regime shift premia

In order to characterise the regime shift premium we present a regime shift model that takes expectations of future shifts to an inflationary regime into account. Such expectations will give rise to a regime shift premium in forward interest rates, which can be seen as a compensation investors demand because they do not view the current price stability objective as fully credible and a switch to a higher inflation level might occur. The size of the regime shift premium depends on the probability of a shift to a high inflation regime assigned by the financial investors. It is likely that such probability assessments in turn depend on the political support for the target, the size and development of the national debt, the degree of central bank independence and the track record of inflation.

Let us first derive an expression for the regime shift component in inflation expectations by decomposing expectations of the future inflation rate, τ years ahead, as

$$\pi^e(\tau) = \pi_0^e(\tau) + \pi_{rs}^e(\tau) \tag{4}$$

where

$\pi_0^e(\tau) = E\,[\pi(\tau) \mid \text{no regime shift}]$ = "normal" expectations about the future inflation rate within the current regime,

$\pi_{rs}^e(\tau)$=the regime shift component of the expected future inflation rate representing the expected change of the inflation rate that regime shifts give rise to.

To obtain a more specific characterisation of the regime shift component, the inflation is assumed to fluctuate around certain levels, π_i, which follow a continuous time Markov chain $\{S(t)\}$ (see Figure 1). There are three possible states:

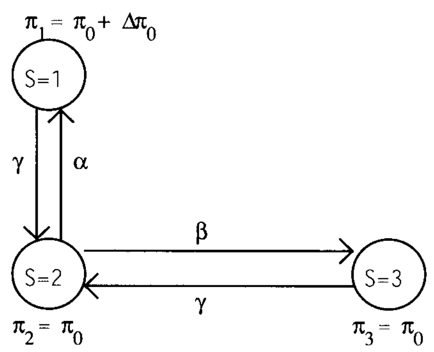
S=1 is characterised by high inflation;

S=2 is characterised by low inflation, but low credibility for the low inflation target;

S=3 is characterised by low inflation and high credibility for the low inflation target

Figure 1

A high/low inflation regime shift model



In the following we will assume that the economy at present is in State 2. This is a critical state, because here the economy can either switch to the high inflation regime (State 1), in which case the inflation rate increases by $\Delta\pi_0$, or to the low inflation regime with high credibility

(State 3). The switching intensities for the states are α and β, respectively, and the switching intensity back to State two from both State one and three is γ. The switching intensity γ is small, if we think that there exists a high degree of persistence of the regimes 1 and 3.[2] The low inflation level, π_0, is supposed to be identical to the declared inflation target.

Let $P(\tau)$ be the probability of being in the high inflation regime τ years ahead conditional on being in State 2 at present. It can be shown that the regime shift component, $\pi_{rs}^e(\tau)$, i.e. the expected increase in the future inflation level, takes the form:

$$\pi_{rs}^e(\tau) = P(\tau)\Delta\pi_0 \tag{5}$$

where[3]

$$P(\tau) = \frac{\alpha}{\lambda}\left[1 - e^{-\lambda\tau}\right] \quad \text{and} \quad \lambda = \alpha+\beta+\gamma$$

The regime shift premium in the forward interest rate (with settlement τ years ahead) is defined, according to the Fisher Hypothesis, as the regime shift component of the inflation expectations:

$$f_{rs}(\tau) = \pi_{rs}^e(\tau) = \varphi(\tau) f_{rs}^{\infty} \tag{6}$$

where $\varphi(\tau) = [1 - e^{-\lambda\tau}]$, and $f_{rs}^{\infty} = \dfrac{\alpha\Delta\pi_0}{\lambda}$

Thus, the regime shift premium is the product of a pure *credibility factor*, f_{rs}^{∞}, and a *credibility sensitivity factor*, $\varphi(\tau)$, showing the credibility effect on the forward rate curve for different horizons, τ. Notice that the credibility sensitivity factor (and hence the regime shift premium) is increasing and concave in τ. These properties as well as the specific functional form will be tested in the next section.[4]

In the model above the credibility (and thus also the regime shift premium) is constant over time. It is, however, possible to extend the set-up above and obtain models in which the credibility is varying stochastically over time. One possibility is to introduce more low inflation states and fluctuations in credibility can then be seen as switches between different low inflation regimes.[5] Another possibility is to let the switching intensities be positive stochastic processes. Since there are strong reasons to suspect that the regime shift premium does vary over time we will in the following extend the set-up above by allowing for a time dependent credibility factor, i.e. $f_{rs}^{\infty} = f_{rs}^{\infty}(t)$. The introduction of a fluctuating credibility factor then provides an explanation of the non-trivial

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[2] The low inflation regime with high credibility (regime 3) is included mainly for preserving the structure from the analysis in Dillén (1996) and Dillén and Lindberg (1997). One may think of regime 3 as an EMU-regime, and the credibility gains for Sweden of joining EMU can be analysed by considering a switch to regime 3. Credibility shocks of this type play an important role in Dillén and Lindberg (1997). In this paper very little is lost by disregarding regime 3 (i.e. $\beta = 0$).

[3] See the Appendix for a derivation of the expression.

[4] Concavity of the regime shift premia is not a general implication of regime shift models of this kind and in a more general setting that takes sticky prices and exchange rate effects into account the regime shift premium can be more complex, see Dillén and Lindberg (1998).

[5] A switch to State 3 can be seen as a positive credibility shock.

phenomenon why long-term interest rates often appear to be more volatile than short-term rates since long-term rates are more sensitive to shocks in the credibility factor, i.e. $\varphi(\tau)$ is increasing in τ.[6]

To find a proxy for the credibility factor, f_{rs}^{∞}, we use the fact that under the assumption that the long-term real interest rate essentially is given by a global real interest rate[7] the difference in long-term forward interest rates between two countries mainly reflects the expected difference in the long-run inflation rate. We assume that Germany has a credible inflation target of 2% (same as the Swedish inflation target) which means that the spread in long-term (10-year) forward interest rate between Sweden and Germany should be an approximation of the credibility factor,[8] $f_{rs}^{\infty}(t)$, i.e.

$$f_{rs}^{\infty}(t) \approx \delta_L(t) \tag{7}$$

Thus, the long-term forward interest rate differential can be seen as a quantitative measure of the degree of credibility of a low inflation policy. Moreover, the long-term forward interest rate differential relative to Germany is judged to be a more precise measure of (imperfect) credibility than the long-term forward interest rate itself, since the former should be less sensitive to international trends in long-term forward interest rates.

The long forward interest rate differential is likely to capture other risk factors than expectations of a shift to an inflationary regime, i.e. default risk or a time varying term premium.[9] In empirical investigations systematic fluctuations in excess returns have often been interpreted as a time varying term premium.[10] However, equation (3) suggests that fluctuations that *seem* to be systematic can also be attributed to fluctuating expectations of future regime shifts. Notice that if regime shifts are rare events it is very difficult (or even impossible if a regime shift does not occur) to statistically distinguish between these two competing explanations due to peso type problems. In other words, the reader is free to consider what the paper calls a regime shift premium as an additional component of the term premium.

Term premia

It is natural to also incorporate a traditional term premium in the analysis. A standard view is that increased variability of interest rates means a higher risk of holding bonds with a maturity exceeding the investment horizon. A measure of risk in this sense is the variability of excess holding returns (the rate of return generated by capital gains during the holding period over the short-term risk free interest rate). To find a proxy for the term premium we first estimate an equation for the excess

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6 Two clarifications should be made about this statement: first, in more elaborated versions of the regime shift model, in which the switching intensities fluctuate, the variability will eventually decline with the term even though the decline will show up for very large τ if γ is close to zero. However, the regime shift model still provides an explanation to why volatility increases (initially) with the term. Second, the variability discussed above is conditioned on that no regime shift occurs. The unconditional variability, which includes the variability that regime shifts cause, is decreasing with the term, as the expectation hypothesis implies.

7 If there are no trends in real exchange rates then this assumption is justified by a real version of the uncovered interest parity.

8 The assumption that the German (implicit) inflation target is two percent is not crucial for the analysis, since the long forward rate interest differential will appear in difference form in the regression.

9 In a fixed exchange rate regime the long-term forward differential should reflect devaluation expectations..

10 See e.g. Fama and Bliss (1987) or Engle, Lilien and Robins (1987) for empirical investigations on US data and Hördahl (1995), Dahlquist and Jonsson (1995) or Sellin (1995) for investigations on Swedish data.

holding return.[11] The generated variance, $\omega(t,\tau)$, gives an estimate of the risk factor measuring the variability of holding returns. We assume that the variance is of GARCH-type[12] and that the forward holding term premia can be written as a constant term, $\rho_0(\tau)$ plus a time-varying term, $\theta(\tau)\omega(t,\tau)$:

$$\rho(t,\tau) = \rho_0(\tau) + \theta(\tau)*\omega(t,\tau) \qquad (8)$$

3. Estimation of regime shift and term premia

We now turn to the estimation of regime and term premia. Equations (3), (6), (7), and (8) suggest the following regression:

$$\Delta\eta(t,\tau) = \mu + \theta(\tau)\Delta\omega(t,\tau) + \varphi(\tau)\,\Delta\delta_L(t) + \varepsilon(t,\tau) \qquad (9)$$

where $E_t[\varepsilon(t,\tau) \mid \text{no regime shifts}] = 0$. To reduce problems of serial correlation (due to the overlapping feature of the time series of forward excess returns) and non-stationarity we take first differences of the time series involved.[13] Weekly data from 9th December 1992 to 24th February 1998 is used. The short-term interest rate is represented by the marginal/repo rate. We assume that there has been no shift to a high inflation regime in this sample.[14] The credibility factor, $\delta_L(t)$, is measured as the ten-year forward interest differential between Germany and Sweden.[15] The variances, $\omega(t,\tau)$, are obtained from a GARCH-M estimation (see the Appendix). The model is estimated for τ equal to 0.25, 0.5 and 1, i.e. for three, six and twelve months. All forward interest rates are continuously compounded and given as effective rates.

From Table 1 it can be seen that the estimated credibility sensitivity factor, φ, is indeed concave and increasing in τ. However, the implied λ value in the case $\tau = 1$ differs from those obtained when the horizon is three or six months even if the difference is not statistically significant. This indicates that the specific functional form of $\varphi(\tau)$ is at best only a reasonable approximation for small τ. The estimate of φ for $\tau = 3$ months is essentially the same as in Dillén (1996), whereas it has decreased somewhat for $\tau = 6$ months.[16] It is notable that the time varying part of the 6-month term

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[11] See the Appendix for estimation of the excess holding return equation. The excess holding return equation is defined as the difference between the holding return during the period $(t,t+s)$ of a zero coupon bond that matures at $t+\tau$ ($\tau > s$), $h(t,t+s)$, and the short-term interest rate, $i(t)$, at time t. Dillén (1996) discusses different estimation techniques based on excess forward return and excess holding return equations. The chosen estimation technique proved to be the most efficient one.

[12] $\omega(t,\tau)$ is assumed to be of GARCH(1,1) type, i.e. $\omega(t,\tau) = \mathrm{Var}_{t-1}[(u(t,\tau)] = \alpha_0(\tau) + \alpha_1(\tau)u^2(t-1,\tau) + \beta_0(\tau)\omega(t-1,\tau)$. $u(t,\tau)$ is the error term from the excess holding return equation.

[13] To test for the appropriate dynamic specification we run regressions of the form: $\eta(t,\tau) = \alpha + \rho(\tau)\eta(t-1,\tau) + \theta(\tau)\omega(t,\tau) + \theta_{-1}(\tau)\omega(t-1,\tau) + \varphi(\tau)\delta_L(t) + \varphi_{-1}(\tau)\delta_L(t-1) + \varepsilon(t,\tau)$. In all regressions ρ is close to unity ($\rho(3) = 0.911$, $\rho(6) = 0.968$, and $\rho(12) = 0.995$) and augmented Dickey-Fuller tests of the time-series for excess forward return indicates that it is hard to reject the hypothesis of a unit root, especially for six and twelve months. Moreover, we also find that $\theta_{-1}(\tau)$ is very close to $-\theta(\tau)$ and that $\varphi_{-1}(\tau)$ was very close to $-\varphi(\tau)$ (the difference is statistically insignificant). The θ and φ-estimates were also very close to the corresponding parameters reported in Table 1. The results of these regression suggest that (9) is an appropriate specification (see Hendry (1995) or Hendry et al. (1984)).

[14] A regime shift in Sweden would probably be an observable event where the current inflation target is given up.

[15] The forward interest rates are estimated by the extended Nelson-Siegel model, see Svensson (1994).

[16] Dillén (1996) did not estimate φ for $\tau = 12$ months.

premium seems to be of less importance in these updated estimations.[17] The high values of R^2 are remarkable in view of the fact that the excess forward return is an unpredictable forecast error under the expectation hypothesis.[18]

Table 1

Estimation of the excess forward return equation:

$$\Delta\eta(t,\tau) = \mu + \theta(\tau)\Delta\omega(t,\tau) + \varphi(\tau)\,\Delta\delta_L(t) + \varepsilon(t,\tau)$$

	$\hat{\mu}$	$\hat{\theta}$	$\hat{\varphi}$	$\hat{\rho}_0$	$\hat{\lambda}$	R^2	DW	SE
τ=3 months	0.0014 (0.131)	0.0439 (7.616)	0.2117 (4.742)	-0.4922	0.09516 [0.530, 1.422]	0.273	2.395	0.176
τ=6 months	0.0023 (0.197)	0.0119 (6.803)	0.3842 (8.124)	-0.6046	0.9696 [0.689, 1.296]	0.382	2.450	1.181
τ=12 months	-0.0014 (-0.106)	0.0015 (6.048)	0.4786 (8.860)	-0.1134	0.6512 [0.0466, 1.296]	0.434	2.261	0.199

Note: $\hat{\rho}_0$ is the implied value of ρ_0 appearing in equation (7) and calculated as the mean of $\eta(t,\tau)$- $\hat{\theta}\,\omega(t,\tau)$- $\hat{\varphi}\,\delta_L(t)$; $\hat{\lambda}$ is the value of λ in the expression $\varphi(\tau)=1-e^{-\lambda\tau}$ that the estimate of φ implies and the 95% confidence interval is given within brackets; t-values within parentheses. DW is the Durbin-Watson statistics for autocorrelation and SE denotes standard error of regression. The number of observations is 259 for 3 months, 246 for 6 months and 220 for 12 months.

As seen from Figures 2a-c the quantitative effects of regime shift premia are, with a few exceptions, much larger than those of the (time varying) term premia. The sizes of the six and twelve months regime shift premia have been above 1%, but they are below 0.5% today. At some occasions, e.g. the summer and fall of 1994, the term premia are substantial.

Interpretation of the results

As emphasised earlier, it is not possible with this estimation technique to formally test for the presence of regime shift premia due to peso type problems. What we have shown empirically is that the long forward interest rate differential relative to Germany appears to be an important factor for explaining systematic movements in excess forward returns. In addition, the impact from the long forward interest rate differential on excess forward return increases concavely with the time horizon. Regime shift models of the type presented in Section 2 are an attempt to provide an explanation for these new and non-trivial findings, but the results can not be seen as proof or strong evidence of the theory. The results can also be interpreted as if we have detected an additional component of a time varying term premium although it is difficult to justify such an explanation from economic theory.

The regime shift premium is also a way of formalising the wide spread use of long (forward) interest rate differential relative to Germany as an indicator of credibility problems. Indeed,

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[17] For τ = 3 and 6 months the constant part of the term premium is negative and rather large in absolute terms. However, these estimates can be shown to be statistically insignificant.

[18] The high values of R^2 do not imply that excess forward returns are unconditionally very predictable, but rather that they are predictable conditional on no regime shift.

Figure 2a

3-month forward rates, regime shift premia and term premia (time varying part) (continuously compounded) (percent)

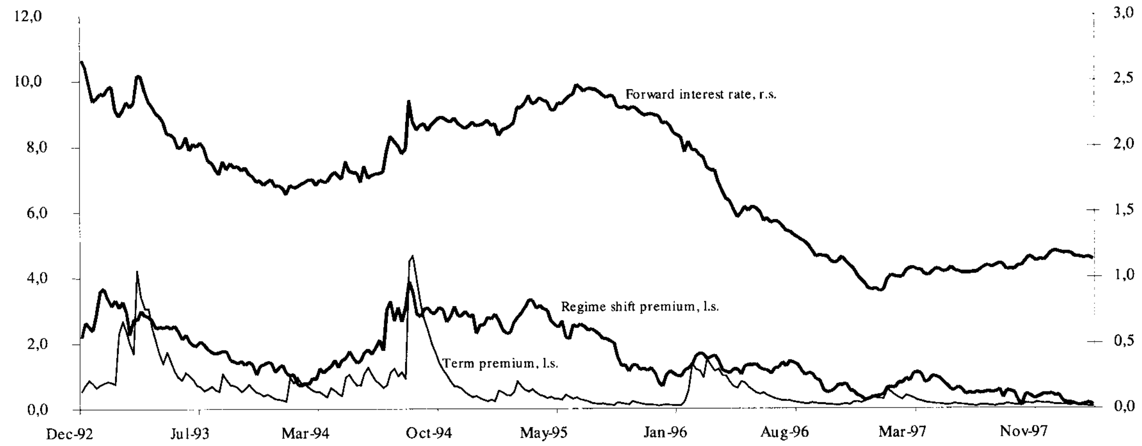


Figure 2b

6-month forward rates, regime shift premia and term premia (time varying part) (continuously compounded) (percent)

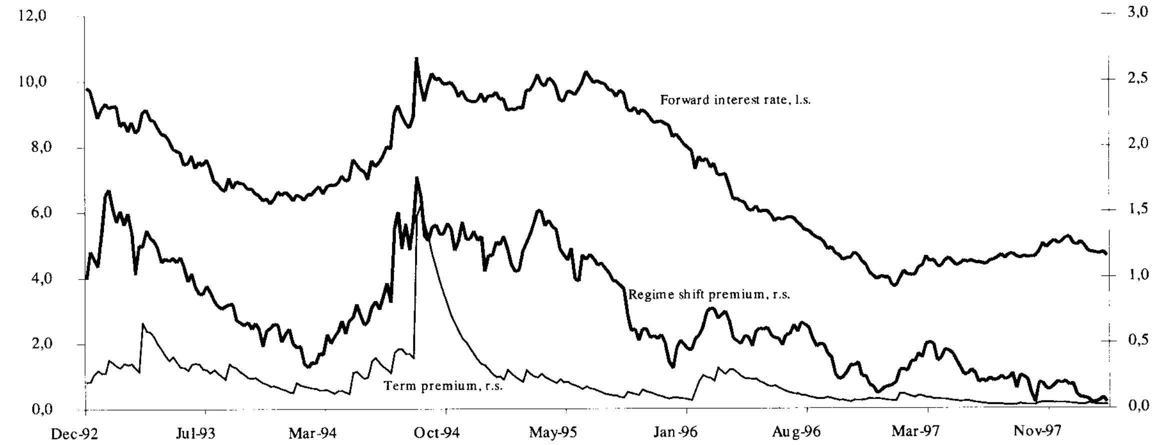
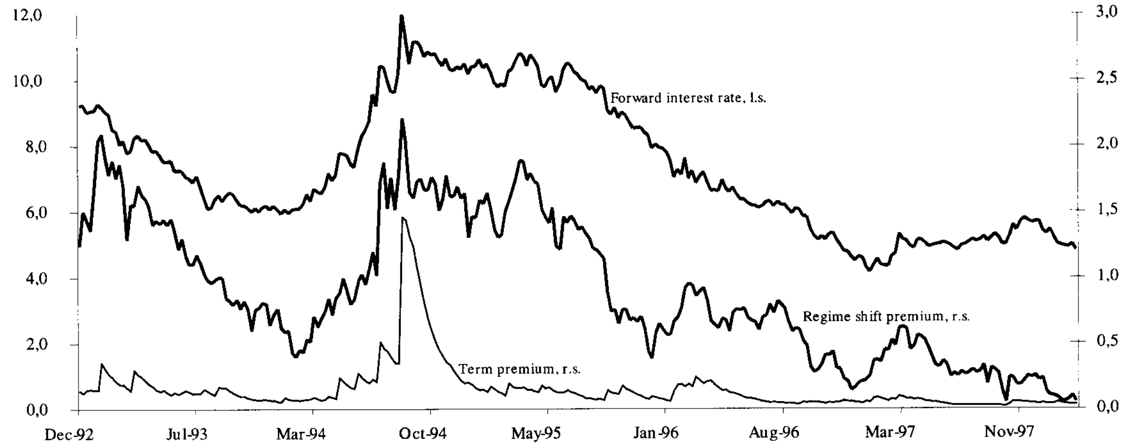


Figure 2c

12-month forward rates, regime shift premia and term premia(time varying part) (continuously compounded) (percent)



The regime shift premium is also a way of formalising the wide spread use of long (forward) interest rate differential relative to Germany as an indicator of credibility problems. Indeed, in periods when credibility problems were obvious in the public debate the regime shift premium has been large. Against this background we will use the term regime shift premium as a measure of the quantitative effects on the forward interest rate curve that credibility problems cause.

Having said that, it is natural to ask what kind of event a regime shift represents. One may argue that a switch to a more inflationary regime initially should imply lower interest rates (in order to generate expansionary conditions in the short run), i.e. negative estimates of $\varphi(\tau)$ for small τ. Indeed, this is a possibility in the regime shift model analysed by Dillén and Lindberg (1998), but a positive regime shift premium is the most natural case also in this model. The Dillén-Lindberg model is an extension of the one presented in Section 2 in that exchange rate effects are also incorporated, and these effects are important to understand why the regime shift premium is likely to be positive for small τ after all. A regime shift would probably be some kind of institutional and political change that will lead to an economic policy that allows for a high and variable inflation rate without necessarily leading to an immediate change of the interest rate policy in an expansionary direction. If a regime shift of this kind occurs, the domestic currency will depreciate drastically when forward looking investors evaluate the implications of the new regime (see Dillén and Lindberg (1998)). Expectations of such a drastic jump in the exchange rate will put upward pressure on short-term interest rates.

4. Investors' expectations of future inflation and monetary policy actions

In this section we will investigate how much information forward interest rates contain about expected inflation obtained from surveys and if investors take possible regime shifts into account when giving their assessment of expected inflation. It is not a priori obvious to what extent survey expectations incorporate regime shifts expectations. If participants in surveys report their inflation expectations, in a mathematical sense inflation expectations should fully reflect regime shift expectations. However, if participants report the most probable outcome of the future inflation then it is likely that regime shift expectations are incorporated to a very limited extent provided that a regime shift is considered to be an unlikely event. In the latter case one should be able to extract more information from forward interest rates by controlling for the regime shift premia. Also, we would expect the interaction between inflation expectations and expectations of monetary policy actions to influence forward interest rates. In order to analyse the effects of regime shift premia and of monetary policy expectations we will consider the following models:

$$\pi^e = \alpha + \beta f(\tau) \quad (10)$$

$$\pi^e = \alpha + \beta_{adj} f_{adj}(\tau) \quad (11)$$

$$\pi^e = \alpha + \beta_{rs} f_{rs}(\tau) \quad (12)$$

$$\pi^e = \alpha + \beta_{adj} f_{adj}(\tau) + \beta_{rs} f_{rs}(\tau) \quad (13)$$

where $f_{adj}(\tau) = f(\tau) - \rho(\tau) - f_{rs}(\tau)$, and $\rho(\tau)$, is the (total) term premium, i.e. $\rho(\tau) = \rho_0(\tau) + \theta(\tau)*\omega(t,\tau)$. The regime shift premium, $f_{rs}(\tau)$, is calculated by taking the estimate for the credibility sensitivity factor, $\varphi(\tau)$ from Table 1 and multiplying this estimate with the long-term forward interest rate differential between Germany and Sweden, $\delta_L(t)$.[19] Since we will only focus on one year forward interest rates ($\tau = 1$) in what follows, the notation for the time horizon, τ, will be dropped.

[19] Since the survey of inflation expectations only comes (approximately) quarterly we have chosen the observations closest in time to Aragon's survey for our estimations. The sample consists of 21 observations, ranging from January 1993 to February 1998.

Equation (10) is the linear forward rate rule analysed by Söderlind (1995). The potential benefit of adjusting forward interest rates for regime shift (and term) premia can be assessed by comparing models (10) and (11). Since we want the real forward interest rate to reflect investors' expectations of the future monetary stance we adjust for the estimated term premium as well. Moreover, the parameter β_{rs} in equations (12) and (13) indicates to what degree regime shift expectations are reflected in investors' inflation expectations. As pointed out by Söderlind (1995), the estimate of β depends on the stochastic properties of the real interest rate.[20] If one of the following two pre-conditions hold: (i) the real interest rate (adjusted for the term premium) is stable or (ii) the real interest rate is negatively correlated with inflation expectations to such degree that this counteracts the estimation effects that a variable real interest rate gives rise to, then we expect: $\beta \approx 1$. Söderlind (1995) found that the forward rate rule (i.e. $\beta \approx 1$) performs fairly well since pre-condition (ii) appears to be fulfilled to a large extent. This finding does not, however, necessarily imply that forward interest rates contain much information about investors' inflation expectations.

However, if investors think that changing prospects of future inflation will lead to a monetary policy response, it is not unreasonable that increasing (decreasing) inflation expectations are associated with increasing (decreasing) expectations of the future short real interest rates, i.e. inflation expectations are positively correlated with the real forward interest rate. In this case only some fraction of a change in the nominal forward interest rate reflects a change in inflation expectations. Thus, we expect: $\beta < 1$.[21] Notice, however, that a β-estimate less than unity might also reflect a high variability of the real interest rate, in which case fluctuations in forward interest rates contain relatively little information about inflation expectations.

Expectations of monetary policy actions

To formalise the idea that investors' expectations of monetary policy are related to their inflation expectations we assume that investors' expectations of the future real interest rate can be written as

$$r^e = r_0 + \phi(\pi^e - \pi_0) + u, \qquad E[u] = 0, \quad Cov[u, \pi^e] = 0 \tag{14}$$

where r_0 is the average expected future real interest rate when investors inflation expectations are on the inflation target (π_0) and u is an error term (uncorrelated with π^e) representing other factors affecting investors' expectations of the future real interest rate. The parameter ϕ can be viewed as how sensitive investors' expectations of the future short-term real interest rate are to changes in investors' inflation expectations. With u set to zero, equation (14) is similar to the instrument rule under inflation targeting (see Svensson (1997)) whereby the real interest rate should increase (decrease) when inflation prospects are above (below) the inflation target. However, inflation prospects in this rule refer to the inflation forecast (conditional on an unchanged interest rate policy) of the central bank which is different from investors' (unconditional) inflation expectations.[22]

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20 See Söderlind (1995) for a further analysis of how the β-estimate depends on the correlation between inflation expectations and the real interest rate.

21 If regime shifts expectations are fully incorporated in investors' inflation expectations then it is reasonable that $\beta_{rs} \approx 1$ since it should be inflation expectations within the regime that mainly affect investors' expectations of the future monetary policy actions.

22 One may ask why investors' expectations deviate from the inflation target. In general investors' inflation expectations will only coincide with the inflation target under the assumptions of (i) strict inflation targeting with no imperfections e.g. model uncertainty, (ii) rational expectations, and (iii) the time horizon of investors' inflation expectations coincide with that of the central bank. In practice these assumptions are not fulfilled. Moreover, investors' inflation expectations often

By using results from single-equation regression analysis it is straightforward to show that the parameter ϕ in equation (14) depends on β and the R^2 from the regressions (10)-(12) in the following way:[23]

$$\phi = \frac{R^2}{\beta} - 1 \tag{15}$$

Thus, forward interest rates may contain much information of investors' inflation expectations even when the β -estimate is small if investors' inflation expectations are strongly linked to expectations about future monetary policy (i.e. future short-term real interest rates) according to (14). One can view formula (15) as a measure of the perceived link between monetary policy and inflation prospects. One implication of (15) is that *expectations of higher future inflation are associated with expectations of higher (lower) future real interest rates if R^2 is larger (lower) than the β-estimate.* It is important to stress that formula (15) is valid only when models are estimated with only one regressor and therefore we apply formula (15) to equations (10), (11) and (12).

The interpretation of the ϕ-estimate in these equations depends on which role regime shift expectations play. It is more natural to think that expectations of future monetary policy are related to normal inflation expectations (within the regime) rather than to expectations about a switch to a high inflation regime, and in the former case the ϕ-estimate in model (11) is of interest. However, it is not unreasonable to assume that investors think that monetary policy will react to increasing regime shift expectations in order to restore credibility. On the other hand, credibility problems may arise due to expectations of a too expansionary monetary policy. Finally, the interpretation of the ϕ is affected by the extent to which expectations of a regime shift are incorporated in inflation expectations obtained from surveys.

Data and graphical inspection

To obtain inflation expectations (π^e) we use Aragon's quarterly survey of financial investors' expectations of average inflation two years ahead for the period 1993q1 to 1998q1.[24] These expectations should normally be a quite good proxy for the expected one-year inflation rate, i.e. the expected rate of consumer price changes one year in the future. We have small sample problems (the data set consists of only 21 observations) and therefore strong quantitative conclusions should be avoided.

The data used in the following analysis is depicted in Figure 3. We see that the unadjusted as well as the adjusted forward interest rate has been more volatile than the inflation expectations. The higher volatility in the forward interest rate arises from the fact that increases in inflation expectations cause the nominal interest rate to rise, the Fisher effect, but also from the fact

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refer to the average rate of change in consumer prices up to a certain point of time (and not the future inflation rate) implying that short-term deviations from the inflation target affect measured inflation expectations.

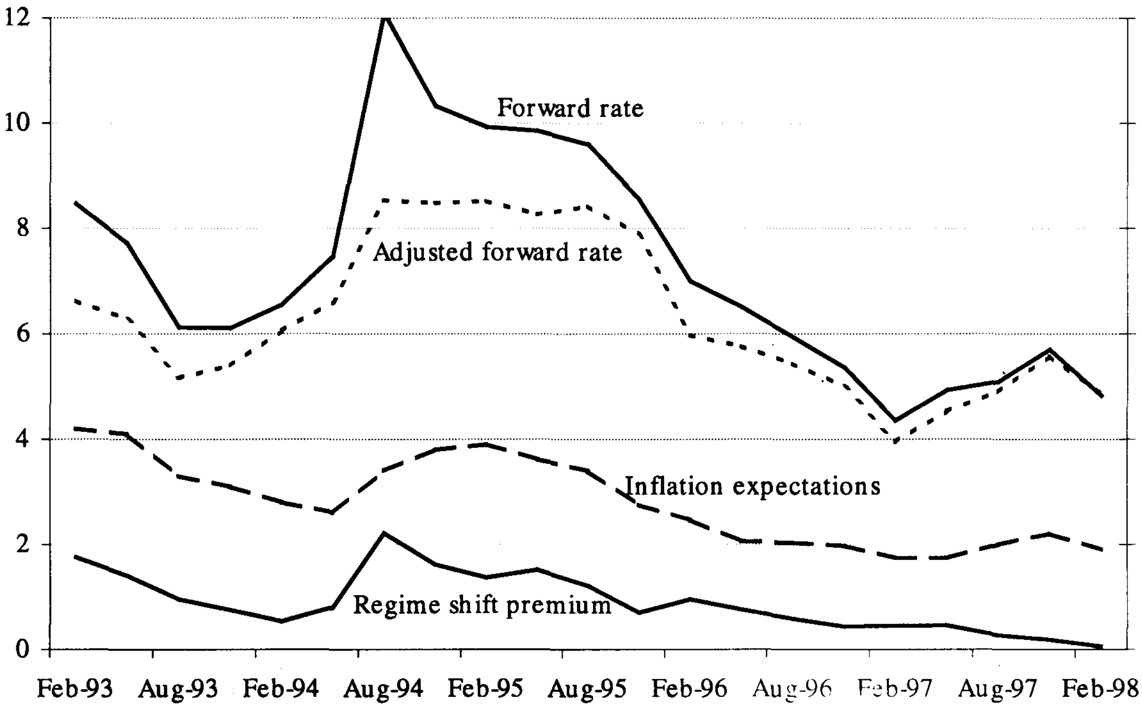
[23] In a regression of the type $y = \alpha + \beta x + \varepsilon$ we have that $R^2 = \hat{\beta}^2 \mathrm{Var}(x)/\mathrm{Var}(y)$ and $\hat{\beta} = \mathrm{Cov}(x,y)/\mathrm{Var}(x)$ implying that $R^2/\hat{\beta} = \mathrm{Cov}(x,y)/\mathrm{Var}(y)$ (i). In this case $(y,x) = (\pi^e, f) = (\pi^e, r^e + \pi^e)$ and (14) implies $\mathrm{Cov}(\pi^e, f) = (1+\phi)\mathrm{Var}(\pi^e)$ (ii). It then follows from (i) and (ii) that $(R^2/\hat{\beta}) - 1 = \phi$. One realises that this expression is valid also when $(y,x) = (\Delta\pi^e, \Delta f)$ and for models (11) and (12).

[24] Aragon Securities Fondkommission AB measures since 1991 every quarter the average expected two-year Swedish inflation of the largest Swedish and foreign investors on the Swedish bond market. We choose not to extend the data to include data from 1991 and 1992 because during this period Sweden had a fixed exchange rate regime and the interest rates were occasionally very high due to devaluation expectations. Our set-up is closely related to credibility problems and monetary policy in an inflation target regime.

that with higher inflation expectations the market anticipates a future tightening of monetary policy, expressed in an increased short-term real interest rate in the future. The latter effect is also incorporated directly in the forward interest rates, which means that their fluctuations will be greater than the fluctuations in inflation expectations. Thus, in contrast to Söderlind (1995), the real interest rate appears to be positively correlated with inflation expectations. The finding that the regime shift premium is positively correlated with investors' inflation expectations may suggest that credibility aspects to some extent are reflected in surveys.

Figure 3

Swedish 12-month forward interest rate and inflation expectations, $\tau = 1$



Sources: Aragon Securities Fondkommission AB and Riksbank.

Moreover, the real one-year forward interest rate (defined as the nominal forward interest rate minus inflation expectations) has fluctuated in the range of 3 to 9%, which is rather high. However, the adjusted real forward interest rate[25] has moved between 2 to 5%, which we think is a reasonable range for the short-term real interest rates in the absence of credibility problems.

Results from regressions

In order to quantify how inflation expectations and forward interest rates interact we estimate equations (10)-(13). All equations are estimated in difference form to reduce problems with serial correlation and non-stationarity.

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[25] The adjusted forward interest rate is also adjusted for the term premium, which normally is included in the real forward interest rate. However, if the real forward interest rate is used as a measure of the expected future real short-term interest rate then the term premium should be excluded. With the exception of a period in 1994 the size of the term premium is small, see Figure 2c.

As seen from Table 2 the adjusted β-estimates ($\hat{\beta}_{adj}$) are approximately in the range of 0.20-0.30 depending on whether the regime shift premium is included or not. The significant estimate of $\hat{\beta}_{rs}$ in estimation 3 suggests that regime shift expectations have an impact on investors' inflation expectations obtained from surveys. However, it is evident from Figure 3 that the regime shift premium is strongly correlated with the adjusted forward interest rate and therefore $\hat{\beta}_{rs}$ comes out insignificant when also the adjusted forward interest rate is included as a regressor. We cannot, however, rule out that the presence of regime shift expectations can explain why inflation expectations of investors tended to be higher than for other groups in 1994-95.[26]

The R^2 is larger when the adjusted forward interest rate is used in comparison with the unadjusted forward interest rate indicating that more information about investors' inflation expectations can be extracted from forward interest rates when they are adjusted for regime shift premia. The presence of a volatile regime shift premium in unadjusted forward interest rates is probably an explanation to why the unadjusted β-estimate ($\hat{\beta}$) is smaller than the adjusted estimate ($\hat{\beta}_{adj}$).

Table 2

Estimates of forward interest rate rules (10) – (13)

1. $\Delta\pi_t^e = \alpha + \beta\Delta f + \varepsilon_t$
2. $\Delta\pi_t^e = \alpha + \beta_{adj}\Delta f_{adj} + \varepsilon_t$
3. $\Delta\pi_t^e = \alpha + \beta_{rs}\Delta f_{rs} + \varepsilon_t$
4. $\Delta\pi_t^e = \alpha + \beta_{adj}\Delta f_{adj} + \beta_{rs}\Delta f_{rs} + \varepsilon_t$

	$\hat{\alpha}$	$\hat{\beta}$	$\hat{\beta}_{adj}$	$\hat{\beta}_{rs}$	$\hat{\phi}$	R^2 {R^2 adj}
1	-0.0855 (0.2321) [0.0691]	0.1617 (0.0003) [0.0364]			1.364	0.382 {0.348}
2	-0.0893 (0.2208) [0.0704]		0.2897 (0.0010) [0.0741]		0.525	0.442 {0.411}
3	-0.0779 (0.3100) [0.0746]			0.4369 (0.0030) [0.1273]	0.393	0.265 {0.224}
4	-0.0749 (0.2255) [0.0596]		0.2365 (0.0067) [0.0766]	0.2251 (0.2236) [0.1782]		0.497 {0.438}

Note: p-values within parentheses, standard errors within brackets. Both are corrected for autocorrelation according to the Newey-West method.

Turning to the interaction between inflation expectations and monetary policy we see that the ϕ estimate of 1.4 when using the unadjusted forward interest rate as regressor suggests that increased inflation expectations lead to expectations of increases in the future short-term real interest rates.[27] The same conclusion holds when the adjusted forward rate is used as a regressor but the

[26] Investors' inflation expectations 2 years ahead were in the range of 3-4% in 1994 and 1995 whereas 1 year inflation expectations of households were between 2 and 3%.

[27] Experiments with Monte Carlo simulations indicate that this estimate as well as the other estimates of ϕ are significant (the estimates deviate from zero by more than three standard deviations).

ϕ estimate is much smaller (0.5). One possible explanation for this result is that the adjustment for a volatile regime shift premium increases the β estimate much more than it increases the informative content (R^2). Moreover, as discussed earlier, the interpretation of the ϕ-estimates depends on the role played by regime shift expectations.[28] In spite of this and the problem of a small sample our results suggest that investors expect future monetary policy to counteract any increases in inflation expectations.

An interesting question is why other studies, e.g. Mishkin (1990) and Söderlind (1995), find evidence indicating a higher β value, closer to unity? One explanation can be that these studies use data that cover periods in which there were less transparent links between inflation prospects and monetary policy actions. The fact that Söderlind (1995) obtained β-estimates greater than the R^2 (indicating a negative ϕ in (14)) supports this view. This study uses data observed in an explicit inflation target regime and it is thus natural that inflation expectations of investors are closely related to expectations of future monetary policy (the real interest rate) and a β-estimate less than 0.5 is in accordance with this.

A general dynamic specification

The models considered so far (i.e. the difference form of equations (10)-(13)) are more or less variants of traditional models in which the Fisher hypothesis can be addressed, and our results are comparable with results from other studies.[29] These models are probably misspecified from a statistical perspective, but the focus of the analysis has been on simple forward rate rules, i.e. to what extent do movements in (forward) interest rates reflect changes in inflation expectations.[30] In order to reduce potential problems of misspecification we extend the previous analysis to analyse the following more general models:

$$\pi^e(t) = \alpha + \delta\pi^e(t\text{-}1) + \beta f(t) + \gamma f(t\text{-}1) \tag{16}$$

$$\pi^e(t) = \alpha + \delta\pi^e(t\text{-}1) + \beta_{adj} f_{adj}(t) + \gamma_{adj} f_{adj}(t\text{-}1) \tag{17}$$

$$\pi^e(t) = \alpha + \delta\pi^e(t\text{-}1) + \beta_{rs} f_{rs}(t) + \gamma_{rs} f_{rs}(t\text{-}1) \tag{18}$$

$$\pi^e(t) = \alpha + \delta\pi^e(t\text{-}1) + \beta_{adj} f_{adj}(t) + \gamma_{adj} f_{adj}(t\text{-}1) + \beta_{rs} f_{rs}(t) + \gamma_{rs} f_{rs}(t\text{-}1) \tag{19}$$

Notice that the difference form models (10) – (13) are obtained if we impose the restrictions $\delta = 1$, $\gamma = -\beta$. The specification in levels corresponds to the restrictions $\delta = \gamma = 0$.

In Table 3 the estimates of the general dynamic specifications are reported. The β-estimates associated with adjusted as well as unadjusted contemporaneous forward interest rates do not deviate much from corresponding estimates in the previous analysis even though $\hat{\beta}_{adj}$ has dropped from 0.24 to 0.18 in the regression including both the adjusted forward interest rates and the regime shift premium. Moreover, we think that this regression generates the most reliable estimates

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[28] The negative ϕ-estimate in the regression that includes only the regime shift premium does not imply a negative correlation between the regime shift premia and inflation expectations since $\mathrm{Cov}(\pi^e, f_{rs}) = (1+\phi)\mathrm{Var}(\pi^e) > 0$ (see footnote 23).

[29] Another reason to use simple models (with only one regressor) is that the interaction between inflation expectations and expectations of the future real interest rates easily can be analysed using equation (15).

[30] A correctly specified model may lead the conclusion that forward interest rates contain very little (if any) additional information about investors' inflation expectations that is not reflected in other regressors. This does not mean that forward interest rates do not contain information about investors' inflation expectations. Forward interest rates can therefore be simple useful indicators for detecting changes in investors' inflation expectations.

since all regressors, with one exception, appear to be significant, whereas the other equations therefore are subject to misspecification to a larger degree. The observation that $\hat{\gamma}_{adj} \approx -\hat{\beta}_{adj}$ in the preferred regression indicates adjusted forward interest rates only reflect information of short-run changes of investors' inflation expectations.

Against this background, our interpretations of the results are that it is advisable to decompose forward interest rates into a regime shift component and an adjusted component. In addition, $\hat{\beta}_{adj}$ should not be very far from 0.2, but a more cautious judgement is that this estimate should be in the range of 0.15 to 0.30.

Table 3

Estimates of forward interest rates rules (16) – (19)

1. $\pi^e(t) = \alpha + \delta\pi^e(t-1) + \beta f(t) + \gamma f(t-1) + \varepsilon(t)$
2. $\pi^e(t) = \alpha + \delta\pi^e(t-1) + \beta_{adj} f_{adj}(t) + \gamma_{adj} f_{adj}(t-1) + \varepsilon(t)$
3. $\pi^e(t) = \alpha + \delta\pi^e(t-1) + \beta_{rs} f_{rs}(t) + \gamma_{rs} f_{rs}(t-1) + \varepsilon(t)$
4. $\pi^e(t) = \alpha + \delta\pi^e(t-1) + \beta_{adj} f_{adj}(t) + \gamma_{adj} f_{adj}(t-1) + \beta_{rs} f_{rs}(t) + \gamma_{rs} f_{rs}(t-1) + \varepsilon(t)$

	$\hat{\alpha}$	$\hat{\delta}$	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\beta}_{adj}$	$\hat{\gamma}_{adj}$	$\hat{\beta}_{rs}$	$\hat{\gamma}_{rs}$	R^2 {R^2 adj}
1	-0.0798 (0.7118) [0.2122]	0.6970 (0.0000) [0.1149]	0.1702 (0.0000) [0.0193]	-0.0534 (0.4732) [0.0727]					0.920 {0.905}
2	-0.1123 (0.6741) [0.2622]	0.7636 (0.0000) [0.0924]			0.3004 (0.0000) [0.0545]	-0.1905 (0.0032) [0.0551]			0.916 {0.900}
3	0.5803 (0.0207) [0.2260]	0.5331 (0.0000) [0.0859]					0.5492 (0.0000) [0.0743]	0.1758 (0.1839) [0.1266]	0.927 {0.914}
4	0.5114 (0.0597) [0.2496]	0.5886 (0.0000) [0.0733]			0.1799 (0.0138) [0.0639]	-0.1792 (0.0004) [0.0386]	0.3686 (0.0371) [0.1600]	0.2574 (0.1454) [0.1669]	0.954 {0.937}

Note: p-values within parentheses, standard errors within brackets. Both are corrected for autocorrelation according to the Newey-West method.

The most important implication of extending the analysis to a more general dynamic specification is that the role of the regime shift premium has increased. The contemporaneous regime shift premium has a significant effect regardless of whether an unadjusted forward interest rate is included in the regression or not. Moreover, the long-run effect of a change in the regime shift component in the preferred regression is around 1.5.[31] If we ignore the insignificant impact from the lagged regime shift premium, the long-run effect is around 0.9 and not statistically significant from 1. We cannot exclude the possibility that $\beta_{rs} \approx 1$, i.e. that changes in the regime shift premium are associated with corresponding changes in investors' inflation expectations on a one-to-one basis. This observation suggests that the (long-run) decline of investors' inflation expectations from about 4% in 1993 to 2% in recent years mainly reflects a corresponding decline of the regime shift premium from 2 to 0%. An inspection of Figure 3 supports this view.

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[31] The long-run effect is calculated as $(\hat{\beta}_{rs} + \hat{\gamma}_{rs})/(1-\hat{\delta})$.

To summarise, we find that the simple forward rate rules in difference form (see Table 2) appear to reflect short-run changes in investors' inflation expectations *within the regime* fairly well, but that these rules are misspecified in the sense that they fail to accurately capture the long-run trend of improved credibility of the low inflation policy (quantified as a decline in the regime shift premium) in the sample period.

Summary and conclusions

In this paper two arguments are put forward to explain the relatively high and volatile development of forward interest rates in Sweden. First, the paper presents a regime shift model, in which investors' fears that the economy will switch to a high inflation regime give rise to a regime shift premium for holding bonds. Estimates of regime shift premia in the forward interest rates are broadly consistent with the implications of the model. Fluctuating regime shift premia are one explanation why forward interest rates have been volatile in Sweden. The analysis also includes term premia, which occasionally are of quantitative importance, but normally appear to be small.

One may argue that the regime shift premium really is a component of a time-varying term premium although it is difficult to justify such an explanation from economic theory. However, there is evidence that investors' inflation expectations obtained from surveys reflect regime shift expectations indicating that the regime shift premium really has something to do with investors' inflation expectations. The analysis suggests that the observed long-run decline of investors' inflation expectations from 4 to 2% can mainly be attributed to gradually improved credibility, manifested by a decreasing (and disappearing) regime shift premium. Regardless of the interpretation of the regime shift premium its quantitative impact on Swedish forward interest rates is hard to ignore.

Another explanation for relatively volatile forward rates is that forward interest rates also reflect investors' expectations about future monetary policy actions (changes in the short-term real interest rate), which tend to amplify the effect on forward interest rates that fluctuating inflation expectations give rise to. An increase (decrease) in investors' inflation expectations is associated with an increase (decrease) in the future short-term real interest rate. This seems to be a new important mechanism that has not been considered in the literature, and an empirical analysis underscores its quantitative relevance. It is likely that this mechanism is more important today than in the eighties, since the increased focus on direct inflation targeting in recent years has created a more transparent link between inflation prospects and monetary policy actions.

From a monetary policy analysis perspective at least two conclusions can be drawn. First, during periods of fluctuating long-term forward interest rates (relative Germany) one should be careful of using short and medium-term forward interest rates as monetary policy indicators. In principle, forward interest rates should be adjusted for regime shifts premia, which reflect long-run trends in credibility for the inflation target rather than cyclical changes of inflation expectations (within the low inflation regime). Notice that this principle is valid even if the quantity that is called a regime shift premium really is an additional component of a time varying term premium. Second, one should only attribute about 20% of the movements in Swedish one-year term forward interest rates (adjusted for the regime shift premium) to changed inflation expectations of investors. Finally, even if the principles mentioned above are important when using forward interest rates as monetary policy indicators, another important aspect highlighted in this study, is the role of inflation expectations in the transmission mechanism. Expectations of higher inflation pressures in the medium term will lead to higher medium term real interest rates before the expected tightening of monetary policy takes place.

Appendix

A. Derivation of expression (7)

Let $Q_i(t)$, i =1, 2 or 3, denote Prob[$S(t+t_0)$ = 1|$S(t_0)$= i]. It follows from the theory of finite state continuous time Markov Chains (see e.g. Karlin and Taylor (1975) p.150-2) that $Q = (Q_1,Q_2,Q_3)'$ satisfies

$$\frac{dQ}{dt} = \Lambda Q, \quad \Lambda = \begin{pmatrix} -\gamma & \gamma & 0 \\ \alpha & -(\alpha+\beta) & \beta \\ 0 & \gamma & -\gamma \end{pmatrix}, \; Q(t_0) = \begin{pmatrix} 1 \\ 0 \\ 0 \end{pmatrix} \tag{A1}$$

It is straightforward to verify that the eigenvalues of Λ are 0, $-\gamma$ and $-(\alpha+\beta+\gamma)$, with corresponding eigenvectors (proportional to) $v_1 = (1,1,1)'$, $v_2 = (\beta,0,-\alpha)'$ and

$v_3 = (-\gamma/(\alpha+\beta),1,-\gamma/(\alpha+\beta))'$, implying that the solution to the system (A1) is of the form

$$Q(t) = av_1 + bv_2e^{-\gamma(t-t_0)} + cv_3e^{-\lambda(t-t_0)}, \quad \lambda = \alpha+\beta+\gamma \tag{A2}$$

The constants a, b, and c can be determined from the initial condition $Q(t_0) = (1,0,0)'$ to be

$$a = \alpha/\lambda, \quad b = 1/(\alpha+\beta), \quad c = -\alpha/\lambda \tag{A3}$$

In particular we have $P(\tau)$ = Prob[$S(t_0+\tau)$ = 1|$S(t_0)$ = 2] = $Q_2(t_0+\tau) = \frac{\alpha}{\lambda}\,[1-e^{-\lambda\tau}]$, which was to be shown.

B. Estimation of excess holding return equation

In this appendix we report the estimates of the excess holding return equation (see Dillén (1996) for details) using the GARCH-M estimation technique (see Engle, Lilien and Robins (1987)). The purpose of this estimation is to obtain an estimate of the variance term ($\omega(t,\tau)$), which is of GARCH(1,1) type and that is used in the excess forward return equation (eq. (5)). As discussed by Dillén (1996), the estimates of the regime shift and term premia (captured by the estimates $\hat{\psi}$ and $\hat{\theta}$) are highly inefficient since the excess holding return equation contains much more noise than the excess forward return equation.[32] The estimates in the variance equation ((eq. (ii)) appears to be more robust. The two-step procedure of using the estimated variance from a GARCH-M estimation in the excess forward return equation (which generates more efficient estimates) appears to be a novelty that generates more robust estimates of parameters.

Estimation of excess holding return equation

(i) $v(t,\tau) = \rho_{0h}(\tau) + \theta_h(\tau)\omega(t,\tau) + \psi(\tau)\delta_L(t) + u(t,\tau), \; E_{t-1}[u(t,\tau)] = 0$

(ii) $\mathrm{Var}_{t-1}[u(t,\tau)] = \omega(t,\tau); \; \omega(t,\tau) = \mathrm{Var}_{t-1}[(u(t, \tau)] = \alpha_0(\tau) + \alpha_1(\tau)u^2(t-1, \tau) + \beta_0(\tau)\omega(t-1, \tau)$

	$\hat{\rho}_h$	$\hat{\theta}_h$	$\hat{\varphi}$	$\hat{\alpha}_0$	$\hat{\alpha}_1$	$\hat{\beta}_0$	$\ln L$
τ=3 months	0.0477 (0.393)	0.0055 (0.193)	0.0411 (0.406)	0.0350 (4.506)	0.2060 (5.652)	0.8086 (14.693)	-246.3
τ=6 months	0.3195 (1.051)	0.0250 (0.929)	-0.1701 (-0.437)	0.0230 (1.254)	0.1083 (5.170)	0.8981 (20.092)	-482.3
τ=12 months	0.9429 (1.146)	0.0066 (0.538)	-0.1733 (-0.1500)	0.9539 (0.9137)	0.1464 (5.080)	0.8572 (19.034)	-717.3

Note: t-values in parentheses; $\ln L$ is the maximised log likelihood value.

[32] The relation between the sensitivity factors ψ and φ is explained in Dillén (1996).

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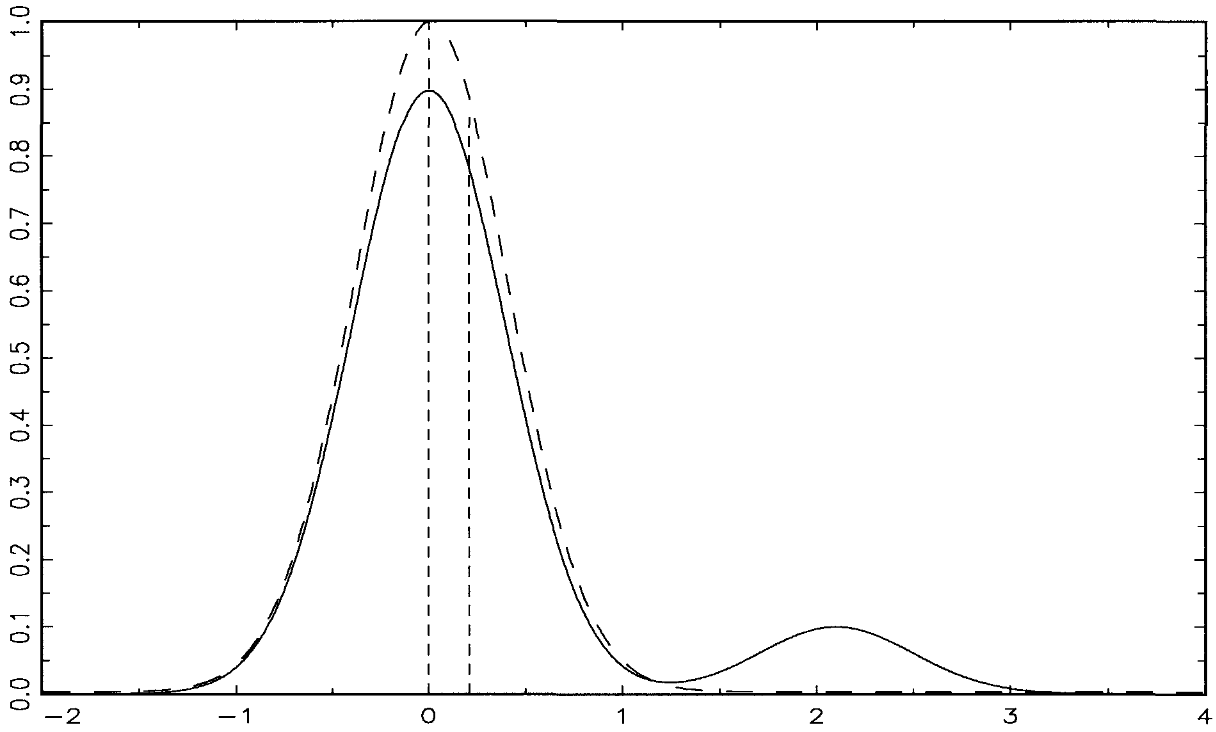
Comments on "Understanding a monetary conditions index" by Hans Dillén and Elisabeth Hopkins

by Pierre Sicsic

This paper deals with the strong forward excess return which has been observed in Sweden at the end of 1994 and in the beginning of 1995 by arguing that the likelihood of an inflationary regime shift increased in this period. The logic of the argument is fine when applied to medium-term rates. It is not as convincing when applied to short-term rates.



In Section 2, the link is made between the forward excess return and the peso problem: if there is some probability that the distribution of interest rates will jump, as, for example, in the plain curve in the above figure, the ex ante expected value of the distribution increases (the mean of the plain curve is above the mean of the dotted distribution). Eventually this unlikely, but large, jump does not occur, and so the observed interest rate remains equal to the mean of the dotted distribution. In Section 3, the spread between Swedish and German forward long-term interest rates is shown to be a proxy of the presence of an inflationary regime shift in Sweden. In the econometrics of Section 4 the (long-term) spread is found to explain the largest part of the forward excess returns on 3 to 12-month rates. At the end of 1994, when short-term rates increased, forward rates increased by a larger amount, and at the same time the long-term forward rate in Sweden increased relatively to the German rate. I do not find that this time correlation suggests that the regime shift in inflationary expectations could imply a peso problem on the short-term rate with a possibility of jump to the right of the distribution. Having in mind the British summer 1992 episode, I would rather suppose that a regime shift would be characterised by a decrease in the short rate. In terms of the figure, the hump on the right should be on the left, lowering the mean of the distribution. Dillén and Hopkins answer this remark by invoking a rise in the short rate in order to prevent a fall of the domestic currency, but such a policy would have signalled that the regime shift had been rejected. Therefore, either one could have expected no regime shift and higher short rates (to convince the market of your intention not to shift), or one could have thought there will be a regime shift and lower short rates. These conjectures point toward larger uncertainty rather than toward higher ex ante expected short rates.

Another path worth following to account for the 1994 forward excess return, which incidentally happened during an election campaign, is therefore the increase in risk, or the variance of the distribution of interest rates. The proxy variable to account for that in the paper might be a poor one, with a single spike in October 1994. Looking at option prices to extract information on the dispersion of the distribution of interest rates should help to settle the questions raised when reading this stimulating paper.

Exchange rate effects and inflation targeting in a small open economy: a stochastic analysis using FPS

Paul Conway, Aaron Drew, Ben Hunt and Alasdair Scott*

Introduction

Specifying the objective of monetary policy as a well-defined target or target range for the rate of inflation is becoming increasingly common among central banks (for example, New Zealand, Canada, the United Kingdom, Sweden, Finland, Australia and Spain). Theory suggests that if the primary cost of inflation arises from consumers' uncertainty regarding the future purchasing power of their incomes, then monetary policy should strive to stabilise a utility-constant consumer price index. In the absence of such ideal indices, central banks have opted to target some available index of consumer prices. For small open economies, movements in the nominal exchange rate often account for a significant part of the variation in these indices via their direct effect on the price of imported goods. This paper examines whether preferable macroeconomic outcomes can be achieved if monetary policy focuses on an index that excludes such direct exchange rate effects on consumer prices.

One argument for targeting indices free of direct exchange rate effects is that the monetary authority should focus primarily on the persistent sources of inflation. As outlined in Mayes and Chapple (1995) and Yates (1995), the design of many inflation targeting regimes includes specific exemptions for disturbances that are expected to result in temporary price level movements only. Depending on how agents form expectations of future inflation, direct exchange rate effects coming through import prices may only result in price level shifts. This arises if agents perceive that a portion of the observed inflation in the CPI index is the result of changes in import prices that are driven by recent movements in the exchange rate and they form their expectations of future CPI inflation by looking through or ignoring these effects.

In Svensson (1997), a model of a small open economy is used to compare CPI inflation and domestic inflation targeting rules. This model characterises direct exchange rate effects on import prices as CPI level effects only. The results from a comparison of policy rules that only attempt to minimise the variance in inflation ("strict" policy rules) suggest that targeting CPI inflation reduces the variance in CPI inflation while increasing the variance in real output, nominal interest rates, the real exchange rate and domestic price inflation in comparison to a domestic price inflation target. The results from a comparison of policy rules that also smooth the variability in real output ("flexible" policy rules) suggest that CPI targeting results in lower variability for most key macro variables except real output and domestic price inflation. The difference between these two types of policy rules arises primarily because they are using different channels to control inflation. The "strict" rules rely on direct exchange rate effects to control inflation and the "flexible" rules work primarily through the output gap channel. The six-, seven- and eight-quarter ahead targeting horizon used in the base-case reaction functions examined in much of this paper means that the monetary authority is working

* Reserve Bank of New Zealand, Economics Department. The views expressed in this paper are those of the authors and may not represent the views of the Reserve Bank of New Zealand. The work presented here has benefited from discussions with Lars Svensson and James Breece. The authors would like to thank the participants at the BIS Model Builder's Meeting for their helpful comments, particularly Francesco Lippi.

primarily through the output gap channel to control inflation. Consequently, these results should be compared to the results under the "flexible" rules examined in Svensson (1997).

In this paper, stochastic simulation experiments with the Reserve Bank of New Zealand's *Forecasting and Policy System* model, FPS,[1] are used to extend, in a number of directions, the work presented in Svensson (1997). First, whereas Svensson's model is designed to represent a generic small open economy, FPS has been calibrated to match the dynamic properties of a specific economy, New Zealand. Second, whereas Svensson's model is linear, the inflation process in FPS is asymmetric in goods market disequilibrium. Third, the way that direct exchange rate effects on import prices enter agents' inflation expectations process is examined, as well as the implications of monetary authority uncertainty about the true expectations process.

The result presented here suggest that targeting domestic inflation relative to CPI inflation reduces the variance in domestic price inflation, output, and the policy instrument; although the variance in CPI inflation is slightly higher, holding all else in the policy reaction function constant. This result holds even if direct exchange rate effects influence inflation expectations and if the monetary authority is uncertain about how direct exchange rate effects influence agents' expectations of inflation. Further, tracing out the CPI inflation/output variability efficient frontier under both CPI inflation and domestic price inflation targeting illustrates that the latter shift the frontier down and to the left. Under domestic price inflation targeting, it is possible to achieve combinations of CPI inflation and output variability that are unambiguously better than those achievable under CPI inflation targeting. Accordingly, for a given variability in CPI inflation, output variability can be reduced by targeting domestic price inflation. Achieving price stability in this way has the ancillary effect of smoothing output. Compared to the results presented in Svensson (1997), these results match those for "strict" policy rules, but not for "flexible" rules which are much closer to the class of policy rules considered here. This difference may reflect the different lag structures in the pass-through of exchange rate effects in the two models or the fact that the results presented here do not consider *optimal* rules. Further work will need to be done to reconcile the differences in these results.

The remainder of the paper is structured as follows. In Section 1, a brief overview of the structure of the FPS model is presented along with the methodology for generating the stochastic disturbances. The stochastic simulation results are presented in Section 2. The final section contains a brief summary and conclusion.

1. FPS at a glance

The Forecasting and Policy System (FPS) model describes the interaction of five economic agents: households, firms, government, a foreign sector, and the monetary authority. The model has a two-tiered structure. The first tier is an underlying steady-state structure that determines the long-run equilibrium to which the model will converge. The second tier is a dynamic adjustment structure that traces out how the economy converges towards that long-run equilibrium.

The long-run equilibrium is characterised by a neoclassical balanced growth path. Along that growth path, consumers maximise utility, firms maximise profits and government achieves exogenously-specified targets for debt and expenditures. The foreign sector trades in goods and assets with the domestic economy. Taken together, the actions of these agents determine expenditure flows that support a set of stock equilibrium conditions that underlie the balanced growth path.

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[1] FPS is a modern macroeconomic model that sits at the heart of the Reserve Bank of New Zealand's new system for generating official economic projections. A complete description of the model and the system can be found in Black, Cassino, Drew, Hansen, Hunt, Rose and Scott (1997).

The dynamic adjustment process overlaid on the equilibrium structure embodies both "expectational" and "intrinsic" dynamics. Expectational dynamics arise through the interaction of exogenous disturbances, policy actions and private agents' expectations. Policy actions are introduced to re-anchor expectations when exogenous disturbances move the economy away from equilibrium. Because policy actions do not immediately re-anchor private expectations, real variables in the economy must follow disequilibrium paths until expectations return to equilibrium. To capture this notion, expectations are modelled as a linear combination of a backward-looking autoregressive process and a forward-looking model-consistent process. Intrinsic dynamics arise because adjustment is costly. The costs of adjustment are modelled using a polynomial (up to fourth order) adjustment cost framework (see Tinsley (1993)). In addition to expectational and intrinsic dynamics, the behaviour of both the monetary and fiscal authorities also contributes to the overall dynamic adjustment process.

On the supply side, FPS is a single good model. That single good is differentiated in its use by a system of relative prices. Overlaid on this system of relative prices is an inflation process. While inflation can potentially arise from many sources in the model, it is fundamentally the difference between the economy's supply capacity and the demand for goods and services that determines inflation in domestic goods prices. Further, the relationship between goods markets disequilibrium and inflation is specified to be asymmetric. Excess demand generates more inflationary pressure than an identical amount of excess supply generates in deflationary pressure.[2] Although direct exchange rate effects have a small impact on domestic prices and, consequently, on expectations,[3] they primarily enter CPI inflation as price level effects.

1.1 Households

There are two types of households in the model: "rule-of-thumb" and "forward-looking". Forward-looking households save, on average, and hold all of the economy's financial assets. Rule-of-thumb households spend all their disposable income each period and hold no assets. The theoretical core of the household sector is the specification of the optimisation problem for forward-looking households. The specification is based on the overlapping generations framework of Yaari (1965), Blanchard (1985), Weil (1989) and Buiter (1989), but in a discrete time form as in Frenkel and Razin (1992) and Black et al. (1994). In this framework, the forward-looking household chooses a path for consumption – and a path for savings – that maximises the expected present value of lifetime utility subject to a budget constraint and a fixed probability of death. This basic equilibrium structure is overlaid with polynomial adjustment costs, the influence of monetary policy, an asset disequilibrium term, and an income-cycle effect.

The population size and age structure is determined by the simplest possible demographic assumptions. We assume that new consumers enter according to a fixed birth rate and that existing consumers exit the economy according to the fixed probability of death. For the supply of labour, we assume that each consumer offers a unit of labour services each period. That is, labour is supplied inelastically with respect to the real wage.

1.2 The representative firm

The formal introduction of a supply side requires us to go beyond the simple endowment economy of the Blanchard et al. framework. The firm is modelled very simply in FPS, but, as with the

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2 Although the body of empirical evidence supporting asymmetry in the inflation process in both New Zealand and elsewhere is growing, the most convincing argument for using asymmetric policy models is the prudence argument present in Laxton, Rose and Tetlow (1994). The evidence for New Zealand is discussed in Black et al. (1997).

3 The direct exchange rate effect on domestic prices is assumed to arise through competitive pressures.

characterisation of the consumer, some extensions are made to capture essential features of the economy. Investment and capital formation are modelled from the perspective of a representative firm. This firm acts to maximise profits subject to the usual accumulation constraints. Firms are assumed to be perfectly competitive, with free entry and exit to markets. Firms produce output, pay wages for labour input, and make rental payments for capital input.[4] The production technology is Cobb-Douglas, with constant returns to scale.

Profit maximisation is sufficient to determine the level of output, the level of employment, and the real wage. FPS extends this framework in a number of directions as firms face adjustment costs for capital and a time-to-build constraint.

1.3 The government

Government has the power to collect taxes, raise debt, make transfer payments, and purchase output. As with households and firms, the structure of the model requires clear objectives for government in the long run. However, whereas households' and firms' objectives arise through explicit maximisation, we directly impose fiscal policy choices for debt and expenditure. The government's binding intertemporal budget constraint is used to solve for the labour income tax rate that supports the fiscal choices. The interactions of debt, spending and taxes create powerful effects throughout the rest of the model; government is non-neutral.

1.4 The foreign sector

The foreign sector is treated as completely exogenous to the domestic economy. It supplies the domestic economy with imported goods and purchases the domestic economy's exports and thus completes the demand side of the model. Further, the foreign sector stands ready to purchase assets from or sell assets to domestic households depending on whether households choose to be net debtors or net creditors relative to the rest of the world. Several key prices affecting the domestic economy are also determined in the foreign sector. The foreign dollar prices of traded goods and the risk-free real interest rate are assumed to be determined in the foreign sector.

1.5 The monetary authority

The monetary authority effectively closes the model by enforcing a nominal anchor. Its behaviour is modelled by a forward-looking reaction function that moves the short-term nominal interest rate in response to projected deviations of inflation from an exogenously specified target rate. Although the reaction function is *ad hoc* in the sense that it is not the solution to a well-defined optimal control problem as in Svensson (1996), its design is not arbitrary. The forward-looking nature of the reaction function respects the lags in the economy between policy actions and their subsequent implications for inflation outcomes. Further, the strength of the policy response to projected deviations in inflation implicitly embodies the notion that the monetary authority is not single minded in its pursuit of the inflation target. Other factors such as the variability of its instrument and of the real economy are also of concern.

FPS is a useful tool for examining the implications of alternative policy reaction functions because agents' expectations are influenced by policy actions. This results from expectations being modelled as a linear combination of a backward-looking autoregressive process and a forward-looking model-consistent process. Modelling expectations in this way partially addresses the critique, initially raised in Lucas (1976), that examining alternative policy actions in reduced-form econometric models gives misleading conclusions. The Lucas critique states that the

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4 We also assume that households own the capital stock.

estimated parameters of such reduced-form models are dependent on the policy regimes in place over the estimation period. Consequently, simulating reduced-form models in which behaviour is invariant to policy actions produces misleading policy conclusions. Although FPS has partially addressed the Lucas critique, a more explicit modelling of agents' learning behaviour would be required to fully address it.

1.6 Stochastic simulations with FPS

Running stochastic simulations with a calibrated model is not as straightforward as with an estimated model. For an estimated model, the properties of the residuals from the estimated equations can be used to pin down the distributions for the shocks that are randomly generated. No such residuals exist for a calibrated model. To generate the shock terms used for the stochastic simulations of FPS we follow a procedure similar to that used in Black, Macklem and Rose (1997). Essentially, the impulse response functions (IRFs) from an estimated VAR are used to calculate the paths for the shocks appearing in the calibrated model's equations. (A more detailed discussion of the methodology can be found in Appendix 1.)

This approach has attractive features as well as weaknesses. First, the VAR itself is a reasonably general representation of the economy and as such captures most of the key temporary disturbances. The VAR approach also leads to shock terms that capture both the serial and cross correlations in the data. The shocks are not interpreted as deep structural shocks, but as summary measures of all the deep structural disturbances that impact the economy at a micro level and, consequently, they should not be expected to be white noise. However, because of the limitations of the New Zealand data, we were unable to estimate a reasonable VAR that include a measure of the supply side. Consequently our application of the VAR technique captures only temporary disturbances. Further, we treat the impulses as if they contain only exogenous disturbances to the economy over the first four quarters. However, even over this horizon the impulses may be capturing some effects from the historical response of policy.

One metric for measuring how well the approach is capturing the stochastic behaviour of the New Zealand economy is to compare the simulated moments from the model with the historical experience. However, given the limitation of New Zealand data one should be cautious about expecting these moments to match very closely. In fact, given that FPS has been calibrated to try to look through the effects that considerable structural change has had on the time-series properties of the data, one might even be dismayed if the moments matched too closely. Examining the moments from 100 draws[5] using the base-case FPS reaction function indicates that the VAR methodology yields results broadly consistent with New Zealand's historical experience. In Table 1, the standard deviations for key macro variables under the base-case FPS reaction function are compared to the historical experience over two sample periods, 1985 to 1997 and 1988 to 1997. The latter period corresponds to the period of inflation targeting. Year-over-year CPI inflation is denoted by π^{cpi}, real output is denoted by y, the nominal short-term interest rate is denoted by rs, and z denotes the real exchange rate. The model generated moments are presented as standard deviations about their equilibrium values. Historical inflation and the nominal interest rate are presented as standard deviations. The real exchange rate is presented as the standard deviation around a linear time trend and the real output standard deviation is calculated relative to potential output.[6]

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5 In order to determine the appropriate numbers of draws, we examined the behaviour of the model's moments as the number of draws were increased. Under a range of policy rules, the results illustrated that the moments did not stabilise until the number of draws reached 70 to 80. Consequently, we choose 100 draws to ensure that the moments were stable enough to allow for sensible comparison.

6 The historical measure of potential output comes from a multivariate filtering technique. The standard Hodrick-Prescott filter is augmented with conditioning information from a Phillips curve relationship, an Okun's law relationship and a

Table 1

Standard deviations

	π^{cpi}	y	rs	z
Model generated moments				
Base-case response	1.1	2.9	4.1	5.0
Milder policy response	1.5	2.7	3.0	4.8
Historical experience				
1985-97	3.9	1.7	5.7	5.0
1988-97	1.7	1.8	3.3	3.5

Under the base-case FPS reaction function, output variability is higher and inflation variability is lower. This suggests that the base-case rule targets inflation more strictly than has been the case historically. Re-running the 100 draw experiment with a milder policy response to projected deviations of inflation from control, produces variability in inflation that is closer to the historical experience. However, model real output variability remains higher than the historical variability.

Ideally, one would like to use the moments that the model would generate under a policy rule identical to that actually followed historically. However, this is not actually feasible given that policy was probably conducted under several different policy rules from 1985 to 1997. Using a rule with a milder policy response and considering the 1988 to 1997 inflation targeting period is one simple attempt to try to more accurately reflect actual historical policy. Clearly more work could be done on the characterisation of historical policy to further improve the degree of comfort with the technique.

2. Targeting domestic inflation versus CPI inflation

2.1 Targeting domestic inflation versus CPI inflation under base-case expectations

The base-case version of FPS is structured such that direct exchange rate effects on import prices only affect the level of the CPI. That is, direct exchange rate effects in the CPI do not impact on inflation expectations. It is worth noting that under inflation targeting, all shocks to prices are allowed to be only levels effects in the long run. Over the near term, the distinction is really about the degree of persistence in prices.

In FPS, CPI inflation is built up by adding imported consumption goods price inflation to inflation in domestic prices. Inflation in domestic goods prices is determined according to a Phillips curve relationship:

$$\pi_t = (1-\alpha)B_1(L)\cdot\pi_t + \alpha\cdot\pi_t^e + B_2(L)(y_t - y_t^p) + B_3(L)(y_t - y_t^p)^+ + f(tot) + g(w) + h(ti) \tag{1}$$

where π represents domestic price inflation, π^e expected inflation, y output, y^p represents potential output, α is a coefficient, $B(L)$ denotes a polynomial in the back-shift operator, $(\cdot)^+$ is an annihilation operator (in this case filtering out negative values of the output gap), $f(tot)$ is a function of the terms of trade, $g(w)$ represents a function of the real wage, and $h(ti)$ a function of indirect taxes. In the base-case model, inflation expectations are given by a linear combination of past and model-consistent values of domestic price inflation:

survey measure of capacity utilisation. A complete description of the methodology can be found in Conway and Hunt (1997).

$$\pi_t^e = (1-\gamma)B(L)\cdot\pi_t + \gamma\cdot C(F)\cdot\pi_t \tag{2}$$

where γ is a coefficient and $C(F)$ is a polynomial in the forward-shift operator.

CPI inflation is given by:

$$\pi_t^{cpi} = \pi_t \cdot B(L)\cdot(pc_t / pc_{t-1}) \tag{3}$$

where π_t^{cpi} represents CPI inflation and pc is the consumption price deflator relative to the price of domestically-produced and consumed goods. The consumption price deflator is a linear combination of the prices of domestically-produced consumption goods and imported consumption goods. The latter term includes the direct price effect of movement in the exchange rate.

The base-case version of the model implies that there is little persistence in inflation arising from direct exchange rate effects. Given this structure, we first examine the stochastic behaviour of the model economy under two alternative formulations of the monetary policy reaction function. The standard reaction function can be expressed as:

$$rs_t - rl_t = rs_t^* - rl_t^* + \sum_{i=1}^{j}\theta_i\left(\pi_{t+i}^e - \pi^T\right) \tag{4}$$

where rs and rl are short and long nominal interest rates, respectively; rs^* and rl^* are their equilibrium equivalents; π_{t+i}^e is the monetary authority's forecast of inflation i quarters ahead, and π^T is the policy target.[7] The number of leads, j, and the weights on them, θ_i, are a calibration choice, and π^e can be defined as any one of a variety of inflation measures.

Our aim is to evaluate which measure of inflation should be targeted. To do this we use the stochastic technique briefly outlined above. Five random shocks are drawn in each period and the model is solved drawing new shocks each period for 100 quarters. This process is repeated for 100 draws and the resulting output averaged into summary measures.

In the first stochastic experiment, the standard reaction function is used. The policy instrument responds to the projected deviations of year-over-year CPI inflation six, seven and eight quarters ahead. In the second experiment, the policy instrument responds to the year-over-year inflation in the price of domestically-produced and consumed goods. The reaction function and the rest of the model are otherwise identical in both experiments.

Table 2

CPI inflation versus domestic price inflation, RMSDs for base-case model

	Targets π^{cpi}	Targets π
π	1.50	1.36**
π^{cpi}	1.13	1.15**
y	3.07	2.70**
rs	4.10	3.90**
rs - rl	2.50	2.34**
z	5.20	5.14**

** denotes that the outcome is significantly different than the outcome under CPI inflation targeting at the 1% confidence level.

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7 The terms of the current Policy Targets Agreement, signed between the Governor of the Reserve Bank of New Zealand and the Treasurer, dictates that the Reserve Bank target an inflation *band* of 0 – 3%. In the base-case version of FPS, the policy target is the mid-point of this band, 1.5%.

The root mean squared deviations of key macro variables are presented in Table 2. The results indicate that targeting domestic price inflation reduces the variance in real output, domestic inflation, the nominal short-term interest rate, and the real exchange rate. Slightly higher variance is recorded for CPI inflation. Qualitatively, these results match those presented in Svensson (1997) for "strict" policy rules. However, given that the inflation targeting rule considered here is closer to Svensson's "flexible" policy rules, these results are importantly different.

This result is obtained because, under CPI inflation targeting, the monetary authority's actions result in temporary disturbances to CPI inflation being partially offset by opposite movements in domestic price inflation. To achieve these offsetting movements in domestic price inflation, monetary policy generates greater variability in real output than it does when it looks through those temporary disturbances. Consequently, real output and domestic inflation variability are significantly lower under domestic inflation targeting. Policy instruments are also less variable as policy itself is less activist; some temporary disturbances generate milder policy responses.

2.2 An alternative specification for inflation expectations

The stochastic experiments presented in Section 2.1 raise some interesting and challenging questions about inflation targeting in a small open economy. However, these results are achieved using the base-case version of the model and may therefore not be robust to different specifications. In this section we test whether the same conclusions still hold under an alternative formulation for inflation expectations.

In the base-case version of FPS, inflation expectations are specified as a function of the core price, the domestic absorption deflator at factor cost. This makes some implicit assumptions about the information that private agents have at their disposal. They correctly perceive that some components of CPI inflation are levels effects only. Alternatively, what if private agents faced a signal extraction problem where they were unable (or unwilling) to decompose CPI inflation into its persistent component and level effects? For New Zealand, this alternative assumption may be reasonable since the data is unable to reveal whether or not direct exchange rate effects influence agents' expectation of generalised inflation.[8]

In the context of a discussion about the choice of target variable in an open economy, it seems important to consider this variation. At a more fundamental level, the appraisal of the target variable under an alternative specification for expectations is very much in the spirit of McCallum (1990). There, the effects of a proposed rule are simulated under two different specifications of the basic structural relationships.

By adding another dimension to the experiments, we now have two extra scenarios to consider. The problem is how to characterise the "state of the world" when inflation expectation effects arise from exchange rate movements. We do this by modifying the expression for inflation expectations that feeds into the Phillips curve so that it is a function of CPI inflation:

$$\pi_t^e = (1-\gamma)B(L)\cdot\pi_t^{cpi} + \gamma\cdot C(F)\cdot\pi_t^{cpi} \tag{5}$$

where γ is the same coefficient as before and $C(F)$ is a polynomial in the forward-shift operator. The same stochastic experiments are run over the two inflation targeting regimes and the results summarised in Table 3. (For comparison, the results from the first two experiments are included as well.)

A comparison of the third and fourth columns of Table 3 tells much the same story as was described in Section 2.1. Even when private agents base their expectations on CPI inflation rather

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[8] In Conway and Hunt (1997), both first and second differences of the exchange rate are included as explanatory variables in a standard Phillips curve equation and both are found to be significant.

than domestic inflation, domestic price inflation targeting is superior if the variability of domestic price inflation, monetary instruments and output is a concern. By effectively filtering out some of the shocks hitting the open economy, targeting domestic price inflation results in less activist monetary policy.

Table 3

CPI inflation versus domestic price inflation, RMSDs when exchange rates have levels or expectations effects

	Levels effects from exchange rate movements		Expectations effects from exchange rate movements	
	Targets π^{cpi}	Targets π	Targets π^{cpi}	Targets π
π	1.50	1.36**	1.42	1.32**
π^{cpi}	1.13	1.15**	1.10	1.09
y	3.07	2.70**	2.85	2.60**
rs	4.10	3.90**	3.49	3.40**
rs - rl	2.50	2.34**	2.10	2.00**
z	5.20	5.14**	4.90	4.90

** denotes that the outcome is significantly different than the outcome under CPI inflation targeting at the 1% confidence level.

A surprising result of these simulations comes from a comparison of expectations formation for given inflation targeting regimes (comparing column 1 with 3 and column 2 with 4). One might expect that by making expectations of generalised inflation a function of CPI inflation, the monetary control problem would be made harder, since direct exchange rate effects and external relative consumption price shocks now influence inflation expectations. In fact, whether targeting domestic price inflation or CPI inflation, there is *less* variability in the macro variables when generalised inflation expectations are formed from CPI inflation rather than from domestic price inflation.

These results might appear somewhat counter-intuitive, until one recalls that in a small open economy, the exchange rate is to some degree influenced by the policy instrument. Since CPI-based expectations include the effects of exchange rate movements, this means that the monetary authority now finds it *easier* to sway expectations than before because of the effect of uncovered interest parity in exchange rate dynamics. Through this channel, the monetary authority has additional control over the monetary problem. On average, the relative importance of this channel is greater than the effect of the exchange rate and external price shocks that are hitting the economy.

2.3 Monetary authority with mistaken beliefs

The results in the previous section imply that CPI inflation targeting is the preferred choice only when its variability is the sole concern of the monetary authority. Less variability in most other key other macro variables is achieved if domestic price inflation is targeted instead. This result holds, more strongly, if direct exchange rate effects influence agents' expectations of generalised inflation.

In all of these experiments, the monetary authority is assumed to know the true structure of the economy. As discussed, the monetary authority understands the nature of private agents' expectations formation, and is able to use this knowledge to its advantage when expectations respond to direct exchange rate effects in prices. In this section, the previous simulations are repeated under alternative assumptions about the accuracy of the monetary authority's perception of the formation of inflation expectations. For example, the monetary authority may believe that expectations of generalised inflation are formed on the basis of domestic price inflation when in fact direct exchange rate effects on prices also influence inflation expectations.

In terms of the stochastic experiment, this "mistake" is made each period. The monetary authority sets monetary conditions on the basis of its belief about the nature of the world, and these monetary conditions are then applied to the "true" model.[9] In the next period, the monetary authority sees that the outcome for the previous quarter was not as it had expected. However, in this period a new set of shocks has also hits the economy, and the monetary authority is unable to unbundle the effects of these new shocks and the excessive or insufficient response of monetary policy in the previous period. Hence there is no learning in this experiment – the monetary authority persists with its view of the world, and sets policy accordingly.[10] The results from these experiments are presented in Table 4 (targeting π) and Table 5 (targeting π^{cpi}).

Table 4

RMSDs from monetary misperception experiments when targeting π

Belief:	Exchange rates have levels effects only		Exchange rates have expectations effects	
State of the world:	Actually levels	Actually expectations	Actually levels	Actually expectations
π	1.36**	1.24**	1.49**	1.32**
π^{cpi}	1.15**	1.05**	1.24	1.09
Y	2.70**	2.60**	2.75**	2.60**
Rs	3.90**	3.90**	3.52**	3.40**
Rs - rl	2.34**	2.47**	2.01**	2.00**
Z	5.14**	5.09**	4.96	4.90

** denotes that the outcome is significantly different than the outcome under CPI inflation targeting at the 1% confidence level.

Table 5

RMSDs from monetary misperception experiments when targeting π^{cpi}

Belief:	Exchange rates have levels effects only		Exchange rates have expectations effects	
State of the world:	Actually levels	Actually expectations	Actually levels	Actually expectations
π	1.50	1.34	1.63	1.42
π^{cpi}	1.13	1.02	1.25	1.10
Y	3.07	2.88	3.07	2.85
rs	4.10	3.97	3.74	3.49
rs - rl	2.50	2.56	2.15	2.10
z	5.20	5.12	4.98	4.90

The first regularity that holds is that regardless of whether the monetary authority's perceptions of the world are correct or incorrect, domestic price inflation targeting yields lower variability in the key macro variables, except for CPI inflation. This can be seen by comparing each column in Table 4 with the respective column in Table 5. This illustrates that the results of previous sections are robust to a more realistic specification of information and policy execution.

9 This simulation technique for examining the implications of the monetary authority being uncertain about the true structure of the economy was first used in Laxton, Rose and Tetlow (1994).

10 Of course, four of these cases – when the monetary authority's beliefs are true – have already been discussed.

Within each of the two targeting regimes, there are interesting results. Consistent with the results in Section 2.2, comparing the first with the second and the third with the fourth columns in each table illustrates that the control problem is easier if direct exchange rate effects influence expectations of generalised inflation. Interestingly, comparing the misperceptions experiment extends this result so that it holds *no matter how the monetary authority believes private expectations to be formed.* When private expectations are based on CPI inflation, there is less variability in all key series. This result holds for both domestic price inflation targeting and CPI inflation targeting when the monetary authority misperceives the structure of the inflation expectations process.

This result suggests that regardless of the true structure of the economy, the monetary authority is better to assume that direct exchange rate effects on prices do not influence inflation expectations. Lower variability in inflation and output results, though with higher variability in instruments. The intuition behind this result is simple: if the monetary authority believes expectations to be a function of domestic price inflation, then it perceives that private agents' do not allow exchange rate effects to enter expectations. Hence it perceives that it will have to do most of its work via the output gap channel.[11] Consequently, it achieves lower variability of inflation and output than it expected, but the misperception results in higher variability in the monetary instruments than would have been the case if it knew the true model.

Essentially, the monetary authority perceives the control problem to be harder than it actually is and it responds more vigorously. Ignoring the increase in instrument variability, the outcome of this "policy error" is a reduction in both inflation and output variability. This outcome illustrates that the base-case FPS rule is not efficient in the sense of Taylor (1994). The rule does not deliver the lowest combination of output and inflation variability achievable. To test whether the results presented thus far are a function of the base-case rule not being efficient, the next section compares the efficient policy frontiers under domestic price and CPI inflation targeting.

2.4 Comparing the efficient frontiers

The overall improvement in the variability of most of the key macro variables under domestic price inflation targeting is notably stronger than the results found in Svensson (1997). Although the results are consistent with the "strict" policy rules considered in Svensson, the base-case FPS reaction function is closer to Svensson's "flexible" rules in the sense that it does not attempt to return inflation to control as quickly as possible by working through the direct exchange rate channel. Thus, the results presented above more strongly favour targeting domestic price inflation than do those presented in Svensson (1997).

One possible reason for this difference is that the rules used in Svensson are optimal rules in the sense that they solve a well-specified optimisation problem. As noted previously, the base-case policy reaction function in FPS is not an optimal rule in this sense. If a well-specified loss function existed, such a policy reaction function could be solved for using a simulation/grid search approach. However, in the absence of a well-defined loss function we can only talk about "efficient" policy rules. As outlined in Taylor (1994), efficient policy rules are defined to be those rules that deliver the lowest achievable combinations of inflation and output variance given the structure of the model economy under consideration. In order to examine whether the results presented previously are obtained because the reaction function considered is not an efficient rule, we trace out the efficient frontiers for forward-looking inflation-targeting rules of the class used in FPS under both CPI inflation targeting and domestic price inflation targeting.

To find the efficient frontiers, we use a grid search technique. In the base-case version of FPS, the reaction function adjusts the policy instrument in response to the projected deviations of

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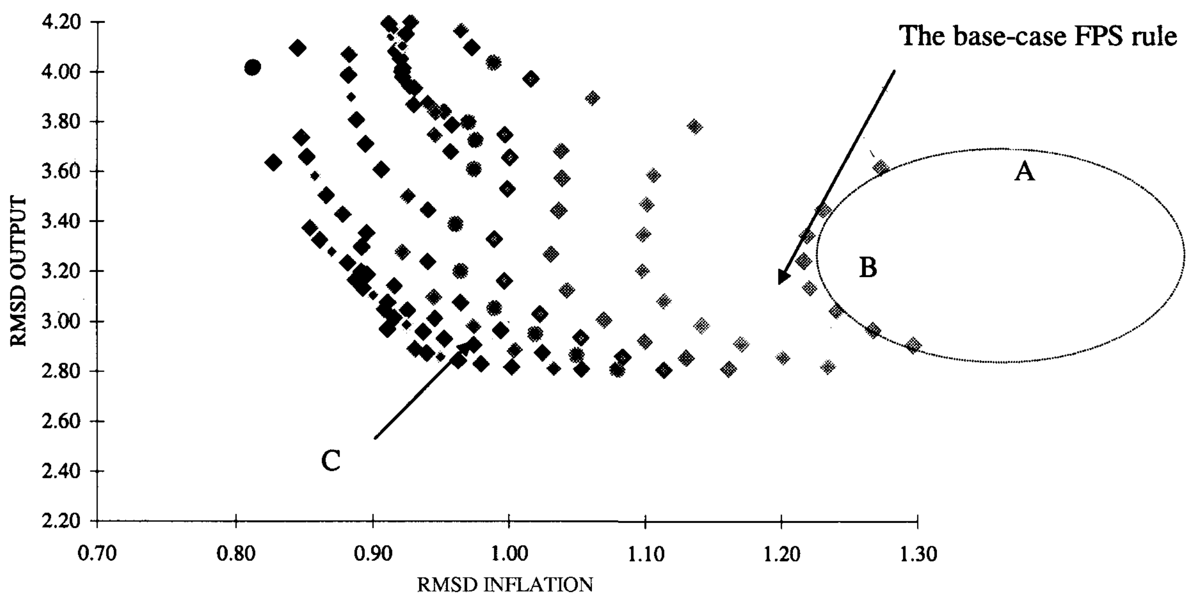
[11] Apart from direct price effects, movements in the real exchange rate will of course shift the trade balance to some extent. This real economy channel will have effects on the output gap, along with interest rates.

inflation from target six, seven and eight quarters ahead. In the base-case version, the weights on the projected deviations of inflation from target are set at 1.4. To determine the set of efficient policy rules, both the magnitudes of the weights and the forward-looking policy horizon are searched over. The forward-looking policy horizon is a three-quarter moving window starting from one quarter ahead and extending to twelve quarters ahead (ten different horizons in all). The weights range from 0.5 to 20. For each rule considered, the resulting properties of the model are calculated by averaging the results from 100 draws, each of which is simulated over a 25-year horizon.

The output/CPI inflation variance pairs under CPI inflation targeting are graphed in Figure 1. The dashed ellipse surrounds the output/inflation trade off that results from holding the weight on the projected deviation of inflation from its target rate fixed at 1 and varying the forward-looking targeting horizon. At point A, the targeting horizon is one, two and three quarters ahead. Moving from point A to point B, the forward-looking horizon is extended to five, six and seven quarters ahead and the variability in both inflation and output are reduced. As the targeting horizon is extended beyond that point, the variability in output is reduced, but only at the expense of increased variability in inflation. For any horizon, increasing the weight up to a point, reduces inflation variability. To reduce both inflation and output variability both the weight and the targeting horizon need to be increased. The results show that the base-case FPS rule lies within the efficient frontier. An efficient outcome is achieved at point C, with a weight of 7 and a targeting horizon of eight, nine and ten quarters ahead. Under this rule, the resulting RMSDs in inflation from target suggest that 90% of the time inflation can be maintained within roughly a 3 percentage point band.[12]

Figure 1

RMSDs in year-over-year CPI inflation and the output gap



In Figure 2, the efficient frontier achievable under domestic price inflation targeting (solid line) is compared to that achievable under CPI inflation targeting (dashed). (A graph illustrating all the outcomes under domestic price inflation targeting can be found in Appendix 2.) The important point here is that the efficient frontier under domestic inflation targeting lies everywhere below the

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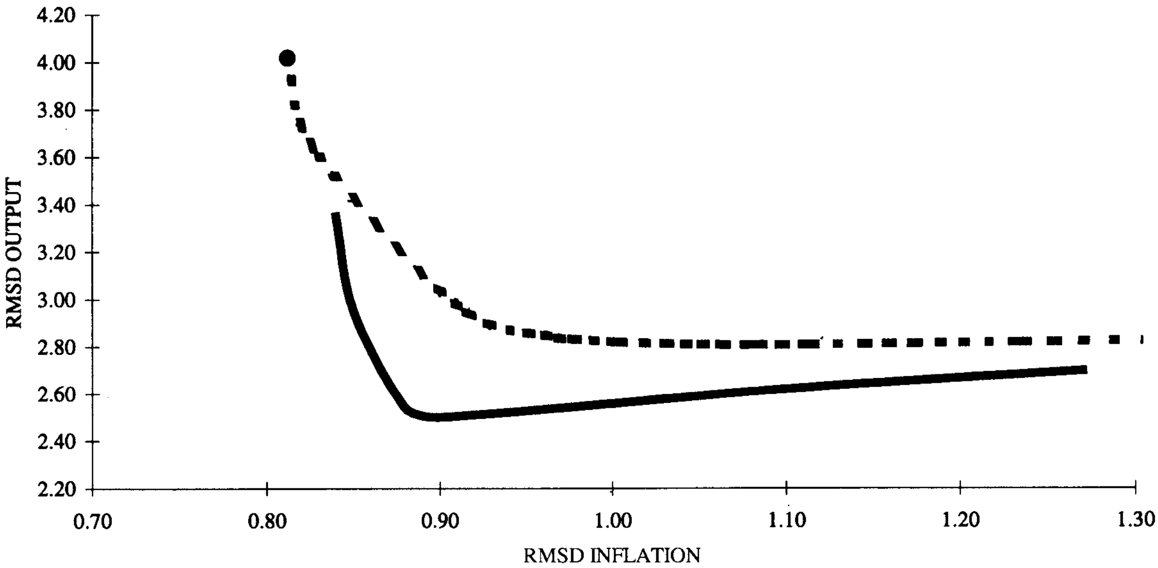
[12] This is calculated as $0.9 \times 1.67 \times 2$. We note that this is very similar to the results found in Turner (1995) for New Zealand.

frontier achievable under CPI inflation targeting. By targeting domestic price inflation, the monetary authority can achieve results that are unambiguously superior to those under CPI inflation targeting.

Figure 2

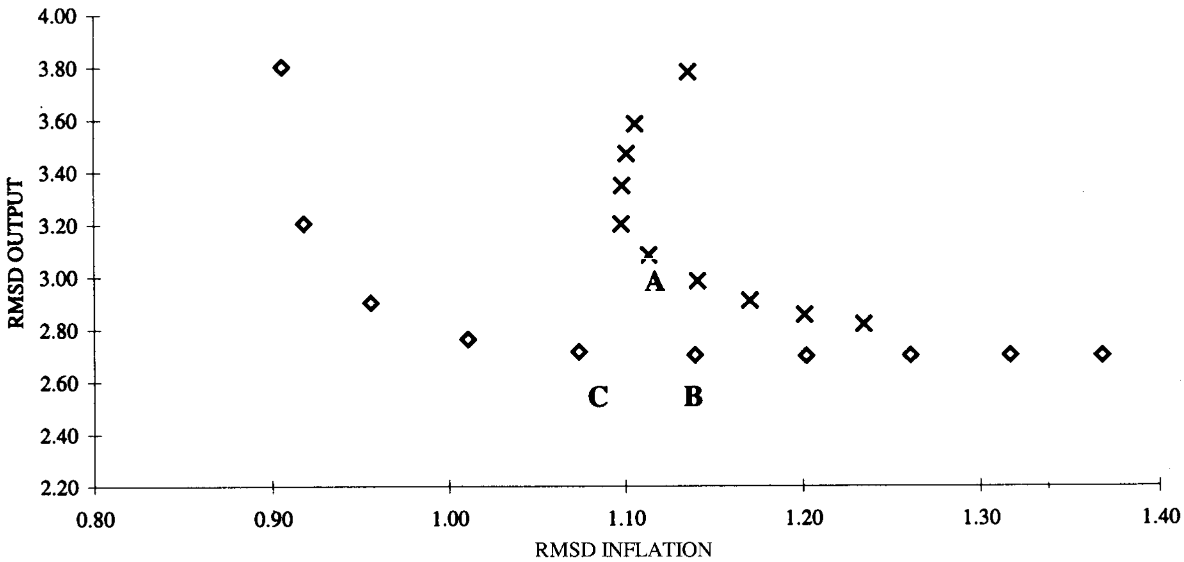
RMSDs in year-over-year CPI inflation and the output gap



Comparing the two targeting regimes under a single weight and varying the targeting horizon illustrates some interesting points. In Figure 3, the results under a rule holding the weight

Figure 3

RMSDs in year-over-year CPI inflation and the output gap holding the weight of 1.4 fixed and varying the targeting horizon



fixed at 1.4 and varying the targeting horizon are compared. The outcomes under CPI inflation targeting are denoted by Xs and those under domestic inflation targeting are denoted by diamonds. The base-case result presented in the paper (i.e., with a targeting horizon of six, seven and eight quarters ahead) can be seen as the move from point A to point B. The relative difference between the two targets are maximised under the short horizon rules and as the horizon gets longer, the results under the CPI targeting rules start to approach those under domestic price inflation targeting rules. This reflects the fact that the longer is the targeting horizon, the smaller will be the impact of direct exchange rate effects on CPI inflation. Moving from point A to point C reduces both the variability of inflation and output. However, instrument variability is slightly higher at point C than it is at point A (RMSD of the nominal interest rate is 4.34 versus 4.17 and the RMSD of the exchange rate is 5.27 versus 5.21). Comparing the impact of shortening the targeting horizon illustrates another interesting point. Under CPI inflation targeting, shortening the targeting horizon from ten, eleven and twelve quarters ahead through to five, six and seven quarters ahead reduces inflation variability but only at the cost of increasing output variability. Under domestic inflation targeting, reducing the targeting horizon in the same fashion reduces both inflation and output variability.

Summary and conclusions

The design of an inflation targeting regime has important implications for the macroeconomic outcomes achieved under these regimes. The particular price index that the central bank strives to stabilise is just one dimension of the design of an inflation targeting regime and this paper has examined one aspect of this issue. Specifically, should central banks in small open economies look through direct exchange rate effects when stabilising inflation? Stochastic simulations of the Reserve Bank of New Zealand's macroeconomic model, FPS, have been used to address the question.

The stochastic simulation results suggest that targeting domestic price inflation reduces the variance in real output, nominal interest rates, the real exchange rate and domestic price inflation with very little increase in CPI inflation variability. Further, the result appears to be robust even if direct exchange rate effects influence agents' expectations of inflation and even if the monetary authority is uncertain about the true expectations process. Tracing out the efficient output/CPI inflation variability frontiers under both CPI inflation and domestic price inflation targeting illustrates that the result is not limited to the base-case FPS reaction function. Targeting domestic price inflation shifts the efficient frontier towards the origin. Under domestic price inflation targeting, the same CPI inflation variability can be achieved with significantly less variability in real output.

Although the robustness tests considered have supported the initial results, more should be done. Although part of the robustness testing traced out the efficient frontiers targeting both CPI inflation and domestic price inflation, the class of policy rules considered may be somewhat restrictive. Future work should be done using an efficient reaction function drawn from a broader range of alternative specifications. In addition, in the stochastic experiments considered there were no permanent shocks. Consequently, there were no permanent movements in the exchange rate. Experiments that allowed for permanent exchange rate movements may yield different conclusions.

Appendix 1: Generating stochastic simulations

This appendix discusses how stochastic simulations were performed using the core structural model in the FPS. First, the VAR model of the New Zealand economy is outlined. Second, the method by which the impulse response functions (*IRF*s) from the VAR are mapped into a set of shocks to the core model equations, such that the model replicates the *IRF*s from the VAR is discussed. The methodology used to implement the shocks stochastically is then presented.

A.1 The estimated VAR model

To capture the stochastic structure of shocks to the New Zealand macro-economy we estimated a six-variable VAR model. The following variables are included in the VAR:

- foreign demand (*fd*)
- terms of trade (*tot*)
- consumption plus investment ($c + i$)
- price level (*cpi*)
- real exchange rate (z)
- slope of the yield curve (*rsl*)

The foreign demand variable is measured as the total industrial production of the OECD. The terms of trade is calculated as the domestic price of exports divided by the price of imports. Shocks to the sum of consumption and investment are interpreted as the result of shocks in aggregate demand. The price index is measured as the consumer price index excluding interest rate effects and GST.[13] The real exchange rate is calculated using the domestic output deflator, the nominal trade weighted index and trade weighted foreign output deflators. Finally, the yield spread is measured as the 90-day paper rate minus the five-year rate. Shocks to this variable are assumed to arise as the result of monetary disturbances induced by the monetary authority. The yield spread enters the VAR in levels and all of the other variables are in log levels.

The variables of the VAR and their associated shocks terms are intended to replicate the stochastic behaviour of macroeconomic disturbances hitting the New Zealand economy. There are, however, a number of omissions. Perhaps the most notable is shocks to the economy's productive capacity. Initially, an estimate of New Zealand's potential output was also included in the VAR. However, given the short length of the sample period there is insufficient stochastic information in the potential output series to produce sensible shock responses. Despite this omission, innovations in the economy's level of productive capacity will in part be captured by the shock terms of the other variables of the system. Stochastic innovations in the domestic price level, for example, can be partially attributed to temporary aggregate supply shocks.

In the reduced form system of equations foreign demand and the terms of trade are modelled as block exogenous on the assumption that New Zealand is a small open economy. Lags of the domestic variables do not, therefore, enter into the equations describing these variables. Foreign demand is assumed to be strictly exogenous in that it is only dependant on its own lags. The equation describing the terms of trade includes its own lags and lags of foreign demand. The equations describing the domestic variables and the real exchange rate are identical and contain lags of all the variables of the system. On the basis of modified likelihood ratio tests, the number of lags in the system is set at four. Ljung-Box Q statistics confirm the lack of serially correlated residuals at the 5%

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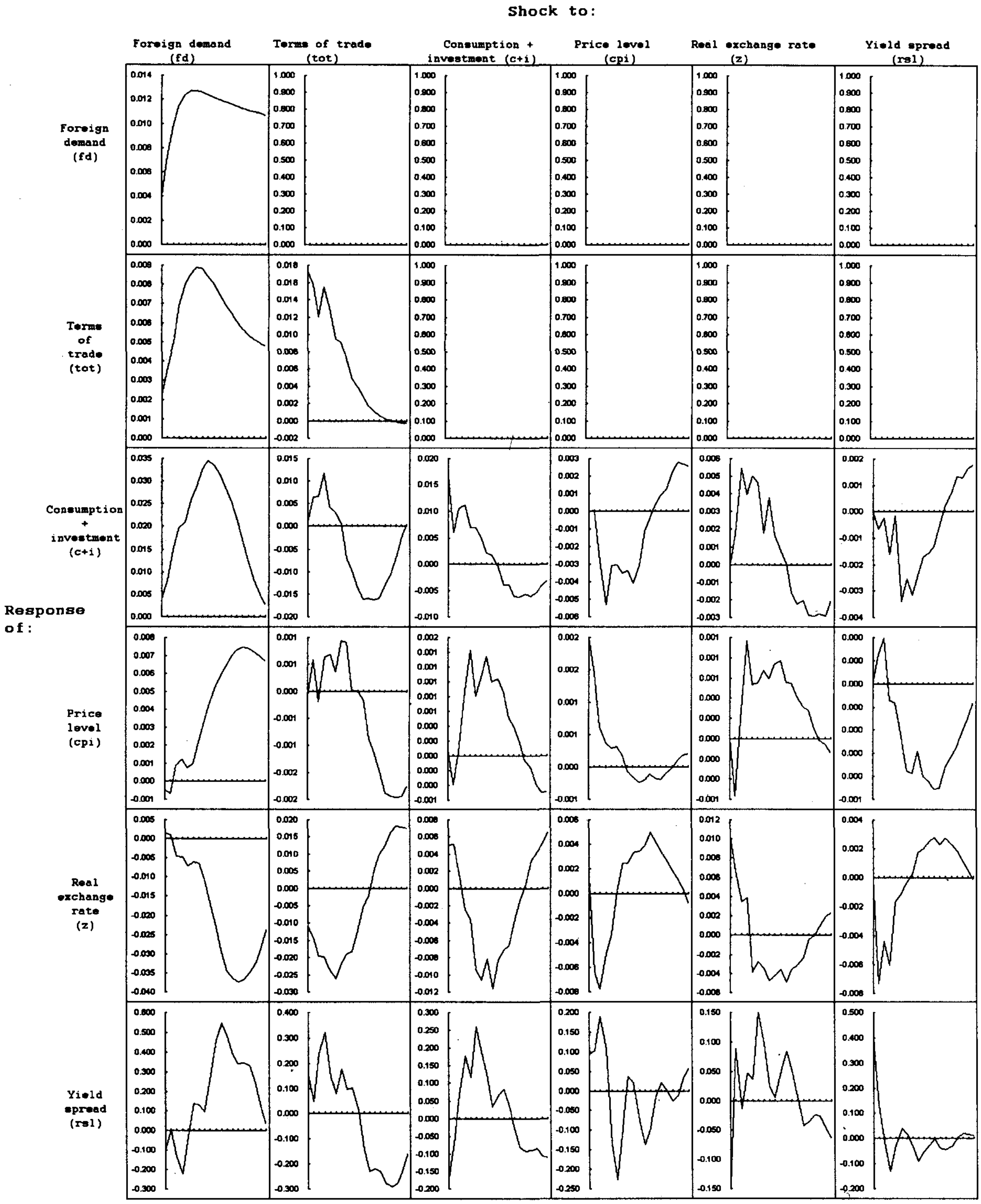
[13] GST is a goods and services tax. This tax was initially implemented in 1986 at 10%. In 1989 GST was increased to 12.5%.

level of significance. The reduced form is estimated over the sample period 1985q2 to 1997q2 using the method of seemingly unrelated regressions.

Figure 4

VAR impulse response functions



To calculate impulse response functions we identify the moving-average representation of the VAR system by imposing a simple contemporaneous causal ordering. The structure of FPS implies an ordering of {*fd*, *tot*, *c* + *i*, *cpi*, *z*, *rsl*}. Foreign demand and the terms of trade are placed causally prior to the domestic variables. This specification extends the assumption of block exogenaity of the foreign sector to the contemporaneous interaction between the variables of the system. In the domestic block, $c + i$ and *cpi* adjust with a lag to shocks in the real exchange rate and monetary conditions. Finally, the monetary authority is assumed to set monetary policy on the basis of contemporaneous (and historical) information.

The impulse responses of the variables in the VAR to each of the six shocks are presented in Figure 4. In all cases the magnitude of the shocks is equal to one standard deviation. The figure should be read vertically; each column shows the response of each variable to a particular shock.

In general, the *IRF*s accord well with the theory of small open economy macro-dynamics. Consider the effect of a one standard deviation shock to foreign demand. The terms of trade improves and the real exchange rate appreciates, consistent with increased demand for New Zealand exports. Aggregate demand, in the form of consumption plus investment, increases inducing a lagged increase in the price level. The monetary authority reacts to increased inflationary pressure by tightening monetary conditions, causing the yield spread to increase. Increased domestic interest rates exacerbates the appreciation of the real exchange rate. In response to a one standard deviation shock to $c + i$, foreign demand and the terms of trade remain unchanged given that they are exogenous. The price level increases after two periods, inducing a tightening in monetary conditions. After three quarters the real exchange rate appreciates.

A.2 Translating VAR impulse response functions into shocks to FPS

In the VAR system, an *IRF* is generated by applying a 1 standard deviation innovation to a variable. The effect of that impulse is seen on all variables in the VAR, subject to the causal ordering of the system and the estimated auto and cross correlations of the variables. Given an impulse, the resultant paths for macroeconomic variables in the VAR can be interpreted as deviations from control or equilibrium.

In the core model, each behavioural equation has an associated shock term. Deviations from control arising from an impulse to the VAR are added to the control levels of those behavioural variables in the core model that most closely match the VAR variables. The core model is then simulated with the behavioural variables concerned exogenous and shock terms on the behavioural equations endogenous. This effectively "backs out" the shocks to the core model that replicate the VAR impulse. This exercise is repeated for each *IRF* produced by the VAR system.[14]

The shocks necessary for the model to replicate the VAR are backed out only for the first four quarters and the policy reaction function in the model is switched off over this period. This is done so that the shock terms are independent of the response of monetary policy over the first year, and consequently independent of the specification of the policy reaction function. By construction, therefore, the methodology assumes that the VAR's *IRF*s are independent of the implicit policy reaction function in the VAR over the first year. Given the long lags between policy actions and real economy responses this is probably a reasonable assumption. Further, it assumes that the response of the policy instrument in the *IRF*s is the result of policy actions alone. This assumption may be a bit strong as the policy instrument may in fact be subject to other innovations in addition to monetary policy actions.

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[14] Two variations of this exercise were examined. In the first, the whole path for the impulses are added to the matching FPS behavioural variables and shock terms are solved for. In the second, the model is simulated quarter-by-quarter with only the contemporaneous effect of the impulse seen each quarter. This makes the problem for the monetary authority harder as the future impact of the impulses are not seen, but is probably more realistic as policy makers do not, in general, know what shocks are hitting the economy at any point in time let alone know the future impact of any shock. As such, this second variation was used in performing the stochastic simulations in this paper.

The shocks required to get FPS to replicate the *IRF*s are serially and cross-correlated. To capture this in such a way that the stochastic simulations can be implemented by drawing random normal (0,1) numbers, the shock terms that appear in the behavioural equations in the model are re-written. A simple example is used below to illustrate how this works.

Example – creating single period shocks that replicate VAR *IRF*s

Suppose that there is only two variables used in the VAR system, $X1$ and $X2$, where an impulse to $X1$ does not effect $X2$ but an impulse to $X2$ effects itself and $X1$. The impulses to the system are mapped to the behavioural *FPS* variables x_1 and x_2 which have associated shock terms x_1_shk and x_2_shk.

Truncated after four quarters, the shocks paths required for the model to replicate the paths arising from the *IRF*s are:

1) Impulse to $X1$

$x_1\_shk = \{\alpha^1_{1,1}, \alpha^1_{1,2}, \alpha^1_{1,3}, \alpha^1_{1,4}, 0,\ldots,0\}$

$x_2\_shk = \{0,\ldots,0\}$

2) Impulse to $X2$

$x_1\_shk = \{\alpha^2_{1,1}, \alpha^2_{1,2}, \alpha^2_{1,3}, \alpha^2_{1,4}, 0,\ldots,0\}$

$x_2\_shk = \{\alpha^2_{2,1}, \alpha^2_{2,2}, \alpha^2_{2,3}, \alpha^2_{2,4}, 0,\ldots,0\}$

where $\alpha^i_{j,t}$ is the numerical solution for the value of the shock term at time t, given the effect the *IRF* i has on the behavioural variable j.

Let ε^i_t be a single period random number at time t. If this random number equals one, then it will generate the shock path required to replicate the *IRF* with the shock structure coded in the behavioural equation as follows:

$x_1\_shk_t = \alpha^1_{1,1}*\varepsilon^1_t + \alpha^1_{1,2}*\varepsilon^1_{t-1} + \alpha^1_{1,3}*\varepsilon^1_{t-2} + \alpha^1_{1,4}*\varepsilon^1_{t-3} + \alpha^2_{2,1}*\varepsilon^2_t + \alpha^2_{2,2}*\varepsilon^2_{t-1}+ \alpha^2_{2,3}*\varepsilon^2_{t-2} + \alpha^2_{2,4}*\varepsilon^2_{t-3}$

$x_2\_shk_t = 0*\varepsilon^1_t + 0*\varepsilon^1_{t-1} + 0*\varepsilon^1_{t-2} + 0*\varepsilon^1_{t-3} +\alpha^2_{2,1}*\varepsilon^2_t + \alpha^2_{2,2}*\varepsilon^2_{t-1} + \alpha^2_{2,3}*\varepsilon^2_{t-2}+ \alpha^2_{2,4}*\varepsilon^2_{t-3}$

Essentially then, the methodology implemented re-writes the shock terms in the behavioural equations to capture all the impulses in the VAR system. The shock paths required to replicate IRF_i for one year will be generated when the random number ε^i_t is takes the value one.

Extending the above example to replicate the VAR in this paper is relatively simple. There are five shock terms that are re-written rather than two. Any individual shock term j appearing in a behavioural equation is represented by:

1) $x\_shk_{j,t} = \sum_{i=1}^{5}\sum_{k=0}^{3}\alpha^i_{j,k+1}\varepsilon^i_{t-k}$

A.3 Generating stochastic simulations

It is convenient to write the full set of behavioural shocks as:

2) $X_t = AE_t$

where: X_t is a vector of the shock terms in the behavioural equations at time t, A is a matrix of the $\alpha^i_{j,t}$ coefficients, E_t is a vector of the ε^i_t random variables that exist from time t-3 to t. This can be de-composed into four sub-vectors e_{t-i}, where each e_{t-i} is a vector of the random numbers at time t-i only.

The stochastic simulations in this paper were generated using the following procedure. First, elements of the vector e_t are drawn from a standard normal (0,1) distribution. Given this vector of random impulses a "shock model" solves for the shock vector X_t. The FPS core model is then simulated with this shock vector exogenous and all the behavioural variables endogenous. This counts as one "iteration" of the model. Typically, in the stochastic simulation experiments considered in this paper, a single "draw" consisted of simulating the model for 100 iterations, where in each iteration a new vector for X_t is generated given the historical and contemporaneous stochastic *iid* impulses in E_t. This exercise is repeated for 100 draws. Furthermore, the drawing of the *iid* random numbers are seeded so that for each set of 100 draws, an identical battery of shocks are generated.

Appendix 2: Efficient frontiers

Figure A 2.1

Domestic price inflation targeting
RMSDs in year-over-year CPI inflation and the output gap

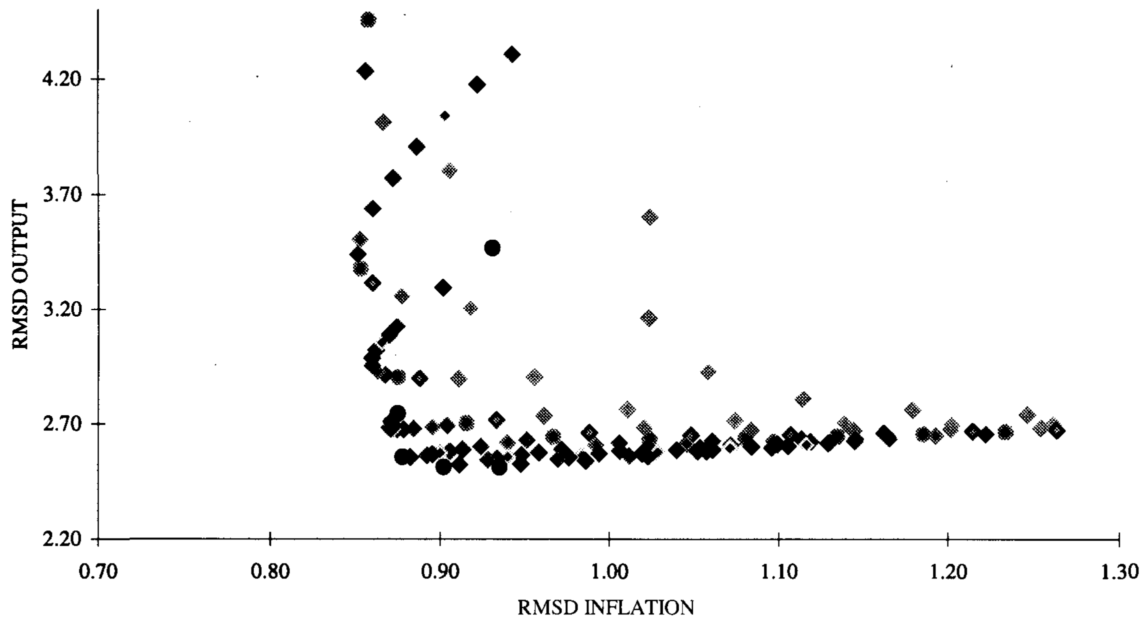
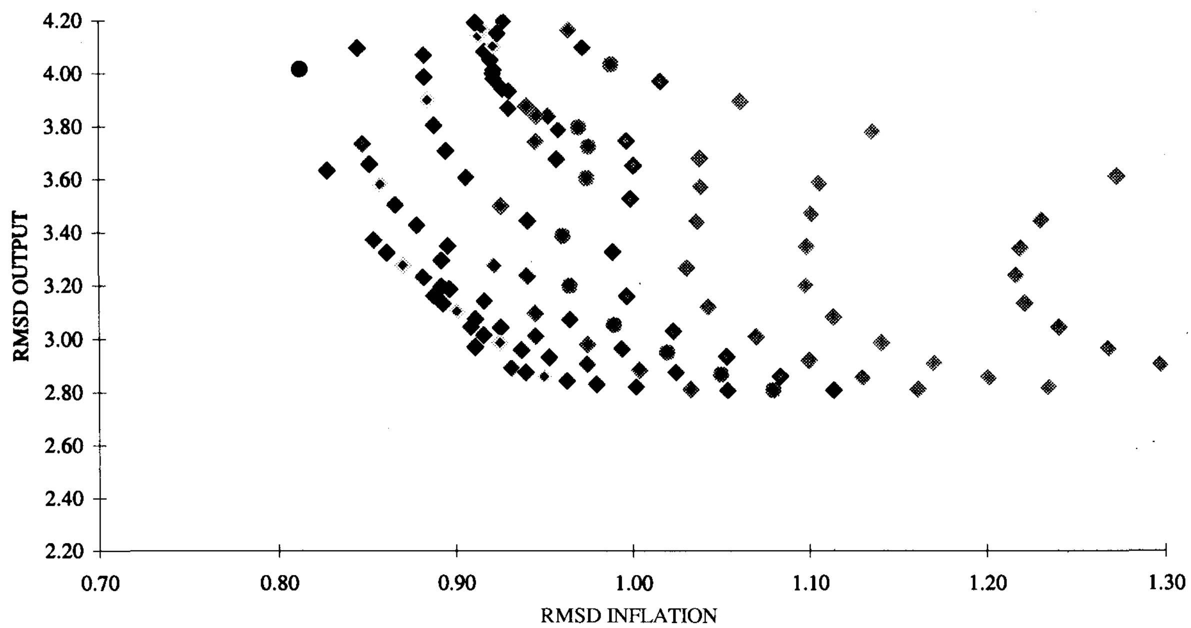


Figure A 2.2

CPI inflation targeting
RMSDs in year-over-year CPI inflation and the output gap



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Comments on "Exchange rate effects and inflation targeting in a small open economy: a stochastic analysis using FPS" by Paul Conway, Aaron Drew, Ben Hunt and Alasdair Scott

by Francesco Lippi*

In the past few years, the diminished reliability of monetary aggregates as an indicator of nominal expenditures has led several countries to adopt, with different degrees of transparency and formalization, a strategy of inflation targeting (*IT*) as a reference framework for monetary policy.[1] The ultimate aim of this strategy is to provide a nominal anchor to inflation expectations, by increasing the transparency and the credibility of monetary policy. Although studies on inflation targets have flourished, their focus has remained mainly theoretical. This paper is therefore very welcome, as it provides us with insights into aspects of *IT* that concern its implementation. Specifically, the paper considers what inflation measure is to be used as a target.

The authors use a macro model of the New Zealand economy (FPS) to assess the differential performance of an *IT* strategy under a CPI inflation target versus a domestic inflation target (i.e. CPI inflation net of exchange rate fluctuations). The "success criterion" adopted by the authors to evaluate performance is given by the variability of a number of macroeconomic variables (real output, nominal interest rates, the real exchange rate), which should ideally be as small as possible. The message that emerges from the paper is that domestic inflation targeting (*DIT*) is preferable to CPI inflation targeting (*CPIT*), because it leads to a lower variability of the main macro variables. The result is robust within a broad class of forward-looking policy rules and under alternative assumptions concerning the expectation formation mechanism. My comments consist of two considerations, regarding:

- the use of FPS in counter-factual simulations
- the *economic* significance of the differences in performance under *DIT* and *CPIT*.

As to the first point, in Table 1 the authors compare the simulated moments (of the main macro variables) generated by FPS with their historic values. They are aware (see Section 1.6) that a "poor" replication of the historic moments by means of a model calibrated to describe policy transmission under a new monetary regime – different from the one that was in place when the historic volatilities were recorded – does not imply that the model is improperly specified. Rather, this could be read as a counter factual simulation pointing at the potential role that *IT* might have had, if it was adopted over that period. If one takes the regime change seriously, it should be expected that under forward-looking inflation targeting, such as the one currently in place, simulated volatilities do not look alike historic ones. I would find it an interesting exercise to investigate what policy rule and expectation mechanism yield simulated volatilities which are in line with the historic ones. However, this question is only marginally addressed in the paper. The authors limit their experimentation to an attempt (only partially successful) to replicate real data by postulating a milder (i.e. less forward looking) policy rule. I would consider a more thorough investigation of this issue a very interesting test of theories of credibility and of the effectiveness of *IT*. Theory suggests that a more transparent and explicitly anti-inflation oriented monetary policy exerts a direct influence on the inflation

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1 For a survey of some recent experiences see Leidermann and Svensson (1995), Haldane (1995) and Bernanke and Mishkin (1997).

expectations formation mechanism, increasing the credibility of policy announcements and making expectations more forward looking. Experimenting with different inflation expectations mechanisms and allowing the objective function of the monetary authorities (implicit in the policy rule) to differ not only with respect to the degree of "conservatism" of alternative targets but also with respect to the targets themselves. Using the empirical model might inform us on how plausible it is to attribute the differences in economic performance to the structural policy changes which have occurred since the early 1990s.[2] Therefore, I think it would be interesting to model the "new rule" jointly with a "new" (for instance less forward looking) inflation expectations mechanism, as the interrelationship between expectation formation and the policy rule (or policy-framework) may presumably be important.

As to the second point, this paper is peculiar in that while most studies of inflation targeting are focused on overcoming the inflationary bias problem of the economy, i.e. on reducing the *average* rate of inflation, the empirical analysis presented here studies the effects that *IT* has on the *variability* of inflation and output. Table 2 shows that under *DIT* the variability of domestic inflation is almost 0.2 percentage points smaller than under *CPIT* (the RMSD are, respectively, 1.36 and 1.50) and the output variability is reduced by approximately 0.4 percentage points (the RMSD are, respectively, 2.70 and 3.07). The authors indicate that the measured variability differential is statistically significant. These results prompt me to address the issue of the *economic*, as opposed to the *statistical*, significance of these numbers: is the performance under the *DIT* regime superior to performance under the *CPIT* strategy, and is the performance (in terms of variability) under either of the inflation targeting regimes superior to the historical one (output and CPI-inflation RMSD amount to, respectively, 3.9 and 1.7 in the 1985-97 period; see Table 1)? Clearly, the basic question concerns the importance of reducing the standard deviation of inflation and output.[3] In my view, the issue is relevant as many economists doubt about the actual welfare benefits of stabilization policy. A back of the envelope calculation can be used to produce a rough estimate of the benefits, in a spirit similar to Lucas (1987). For obvious reasons of space, I will make use of some gross approximations here. Let us concentrate on the welfare effects of output variability, using it as a proxy of consumption volatility (an assumption that is likely to *overemphasize* the benefits of stabilization policy as consumption can be smoothed intertemporally). If we take the utility function of a representative agent to be logarithmic, a 2% reduction in the volatility of consumption corresponds, in terms of welfare units, to an increase of *average* consumption of 0.02%.[4] Thus, even a complete elimination of the *volatility* seems to yield a rather small gain, compared with the effect of a (permanent) change in the growth rate of consumption. Considering that the results reported by the authors indicate only modest volatility differences of the main macro variables under the alternative (simulated) inflation targeting regimes, smaller than the volatility reduction hypothesized in the above example, it seems legitimate to wonder whether the volatility of macroeconomic variables is the right metric to assess the performance of the macroeconomy under alternative *IT* regimes and, even granting that it is, whether the reported differences in volatility associated with the two *IT* regimes considered are relevant. I think this paper, by providing a quantitative assessment of the impact that *IT* may have over

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2 It need not be remembered that the existence of multiple targets is one of the essential preconditions for credibility problems to exist, and that central banks in many countries have shifted away from the multi-target discretionary oriented monetary policies of the seventies only recently.

3 I am deliberately considering *variabilities* instead of *averages* to keep in line with the focus of the paper.

4 The calculation is done by approximating with Taylor expansion the utility function $U[C_t] = \ln[C_t]$ around the steady state consumption, $\overline{C}_t$, and evaluating expected losses ex-ante. This yields:

$E[\ln C_t] = \ln[\overline{C}_t] - \frac{1}{2}\left[\frac{\sigma_c}{\overline{C}_t}\right]^2$, where σ_c is the standard deviation of consumption. The order of magnitude of the benefits arising from reductions of consumption volatility remain "small", compared with changes in the rate of consumption, within a broad class of CRRA utility functions (see Lucas, 1987).

the volatility of macro variables, raises the question of how relevant the volatility improvement that can be obtained really are. My preliminary rule-of-thumb assessment of these gains is that they are not very large; but this is admittedly a very indirect check of the hypothesis (and certainly a partial one, as I completely avoid an assessment of the costs of inflation volatility), to which I hope more research will be dedicated in the future.

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Inflation versus monetary targeting in a P-Star model with rational expectations

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Introduction

In its report on the elements of the future monetary policy strategy of the European System of Central Banks in Stage III of monetary union, the European Monetary Institute rejected the strategies of exchange rate targeting, interest rate targeting and nominal income targeting as inappropriate (EMI (1997)). Inflation targeting and monetary targeting are the remaining possible strategies for future monetary policy in Europe. The choice between these alternative strategies crucially depends on the properties of the monetary transmission mechanism in the common currency area to be created.

In order to investigate how inflation targeting and monetary targeting work and how they relate to the properties of the monetary transmission mechanism, an appropriate theoretical framework is needed. From a purely logical point of view, it seems to be essential to let the money stock play an active role, especially with respect to its influence on the future price level. Numerous models on which the analysis of inflation targeting rests do not meet this requirement.[1] In these models, the money stock is determined only passively on the basis of a money demand function which is given as a recursive element of the respective model. Thus, a monetary policy strategy aimed at controlling the money stock is inefficient *a priori*.

Against this background, the paper presents a stylised model of a small open economy drawing on the P-Star approach which is considered to be a more adequate reference framework. According to the P-Star approach, which is generally used in isolation in the relevant literature, inflation is considered to be a monetary phenomenon in the long run which results from an excessive money supply by the monetary authority. The approach is empirically motivated by the fact that there is a long-term relationship between the money stock and the price level.[2] In view of the importance of forward-looking behaviour on the part of economic agents for the transmission of monetary impulses, forward-looking rational exchange rate and inflation expectations are taken into account.[3]

The alternative monetary policy strategies are implemented within the theoretical model by specifying appropriate feedback rules for monetary policy. According to the realisation of the respective monetary policy target, i.e. the inflation target in the case of inflation targeting and the monetary growth target in the case of monetary targeting, the policy rules determine the nominal short-term interest rate which is regarded as the monetary policy instrument. By endogenising monetary policy a nominal anchor is obtained for the forward-looking expectations of economic

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1 See, for example, Blake and Westaway (1994, 1996a), Britton and Whitley (1997), Fagan and Vega (1997), Fillion and Tetlow (1994), Fuhrer and Moore (1995a, 1995b) and Fuhrer (1997).

2 See the empirical findings in Issing and Tödter (1995) and Tödter and Reimers (1994).

3 Blake and Westaway (1994, 1996a) deal with the role of inflation expectations for the operation of inflation targeting in particular.

agents that are directed towards future monetary policy.[4]

The model of a small open economy with rational expectations based on the P-Star approach is developed first in Section 1. Then, in Section 2, the operation of monetary targeting and inflation targeting is analysed on the basis of two simulation scenarios – a transitory aggregate demand shock and a transitory money demand or velocity shock. Section 3 concludes with a short summary of the theoretical findings. These are discussed subsequently with respect to problems of economic policy which arise when inflation or monetary targeting is to be put into practice. An appendix describes the method employed in solving and simulating the model.

1. The model

In the short run, inflation is determined by monetary as well as by real factors. It is generally agreed, however, that inflation is a purely monetary phenomenon in the long run. Relying on a stylised IS/LM model of a small open economy based on the P-Star approach with rational inflation and exchange rate expectations, these features of monetary transmission will be illustrated below.

1.1 The IS/LM framework

Except for home and foreign interest rates, the model is specified in log-linear form:

$$y_t - y_t^* = \alpha_1 (y_{t-1} - y_{t-1}^*) - \alpha_2 (r_t - r_t^*) + \alpha_3 \left[(e_t + p_t^f - p_t) - (e + p^f - p)_t^* \right] + \varepsilon_t^y \tag{1}$$

$$m_t - p_t = -\beta_1 i_t + \beta_2 y_t + \varepsilon_t^m \tag{2}$$

$$r_t = i_t - E_t [\Delta p_{t+1}] \tag{3}$$

$$\Delta p_t = \gamma_1 E_t [\Delta p_{t+1}] + (1 - \gamma_1) \Delta p_{t-1} + \gamma_2 (p_t^* - p_t) \tag{4}$$

$$e_t = E_t [e_{t+1}] + i_t^f - i_t \tag{5}$$

Here, y denotes real output, r the real interest rate, e ($e + p^f - p$) the nominal (real) exchange rate, p (p^f) the (foreign) price level, m nominal money, i (i^f) the (foreign) nominal interest rate and ε^y (ε^m) aggregate demand (money demand) shocks. Equilibrium values are marked with a star. All parameters are restricted to be non-negative. Furthermore, $\alpha_1, \gamma_1 < 1$ is assumed to hold.

Aggregate demand as well as money demand shocks follow a first order autoregressive process

$$\varepsilon_{t+1}^j = \rho^j \varepsilon_t^j + \eta_{t+1}^j, \qquad |\rho^j| < 1 \text{ for } j = y, m \tag{6}$$

where η^j are serially uncorrelated innovations with expectation zero realised in transition from period t to period t+1.

Finally, Δ denotes the difference operator and $E_t\,[\cdot] \equiv E\,[\cdot\,|\Omega_t]$ the expectation operator conditional on the information set Ω_t available in period t with $\Omega_t \supseteq \Omega_{t-1}$. The information set Ω_t contains the realisations of the exogenous variables and past endogenous variables which are

4 The necessity of endogenising monetary policy to guarantee the determinacy of the model solution in the presence of forward-looking expectations was recently re-emphasised by Fisher (1992), Chapter 6, and by Blake and Westaway (1994, 1996a).

predetermined.[5]

Equation (1) describes the deviation of real output y_t from equilibrium output y_t^*, i.e. the output gap $y_t - y_t^*$, which is determined by aggregate demand in the short run. The output gap depends on the output gap of the last period, on the deviation of the real interest rate r_t from the equilibrium real interest rate r_t^*, on the deviation of the real exchange rate $e_t + p_t^f - p_t$ from the equilibrium real exchange rate $(e + p^f - p)_t^*$, which is determined on the basis of purchasing power parity, and on the aggregate demand shock ε_t^y. In equation (2), real money demand $m_t - p_t$ depends on the nominal interest rate i_t, on real output y_t and on the money demand shock ε_t^m. As equation (3) shows, the real interest rate r_t is determined by means of the Fisher equation as the difference between the nominal interest rate i_t and the expected inflation rate $E_t\,[\Delta p_{t+1}]$.

In equation (4), the current inflation rate Δp_t is specified as a function of the *price gap* $p_t^* - p_t$ which measures the deviation of the price level p_t from the *equilibrium* price level p_t^*. Furthermore, it includes both a *backward-looking* element Δp_{t-1} which reflects the persistence of the inflation process to be observed in reality and a *forward-looking* expectation element $E_t\,[\Delta p_{t+1}]$ which is implicitly directed towards *future* price disequilibria as the cause of future inflation and which has an immediate effect on current inflation.[6]

The evolution of the nominal exchange rate e_t is governed by uncovered interest parity (equation (5)), i.e. arbitrage transactions of international investors lead to the equalisation of expected returns on home and foreign financial assets. The arbitrage transactions induced by the interest rate differential and the expectation of an exchange rate depreciation guarantee a continuous equilibrium in the international financial markets.

The specification of the inflation equation (4) guarantees the compatibility of the model with *any* equilibrium inflation rate.[7] In addition, as the nominal variables are homogeneous in the price level, the neutrality of monetary policy with respect to real variables holds in the long run.[8] The real interest rate and the real exchange rate assume their equilibrium values in the long run and, therefore, output cannot deviate permanently from its equilibrium value, i.e. the natural rate hypothesis holds.

As the factors underlying the real equilibrium values are not specified within the model, these are set equal to zero for the sake of simplicity, i.e. $y_t^* = r_t^* = (e + p^f - p)_t^* = 0$. For the same reason this simplification is carried out for the foreign variables, i.e. $p_t^f = i_t^f = 0$. The equilibrium price level p_t^* and the nominal interest rate i_t remain to be specified.

1.2 The P-Star approach

The determination of the equilibrium price level p_t^* and thus of the price gap $p_t^* - p_t$ is based on the P-Star approach.[9] The starting point is the equation of exchange (in logarithms) solved for p_t:

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5 For the definition of terms see the Appendix.

6 Fuhrer and Moore (1995a), for example, show in a model with a *staggered contracts* (real-) wage equation that a combination of backward-looking and forward-looking elements may generate a high degree of inflation persistence.

7 This is warranted by the restriction that the parameters of the expected future and of the past inflation rate add up to unity.

8 In the Appendix it is shown that the model has an equivalent representation in stationary real levels on account of its homogeneity in the price level.

9 For details see Hallman, Porter and Small (1991), Deutsche Bundesbank (1992), Tödter and Reimers (1994) and Issing and Tödter (1995).

$$p_t = m_t + v_t - y_t \qquad (7)$$

where v_t denotes *velocity*, i.e. the inverse coefficient of liquidity holdings.

The equilibrium price level p_t^* is defined as the price level which, for any given amount of money in circulation, is obtained when velocity and output assume their equilibrium values v_t^* and y_t^*:

$$p_t^* \equiv m_t + v_t^* - y_t \qquad (8)$$

It immediately follows from the preceding equations (7) and (8) that the price gap $p_t^* - p_t$ is composed of the output gap $y_t - y_t^*$ and the *liquidity gap* $v_t^* - v_t$:

$$p_t^* - p_t = (y_t - y_t^*) + (v_t^* - v_t) \qquad (9)$$

This decomposition illustrates that inflationary pressures exist not only when production capacity is excessively utilised but also when velocity is lower or liquidity holdings are higher than in equilibrium, i.e. a monetary overhang exists.

The liquidity gap is unobservable. However, according to Tödter and Reimers (1994) v_t can be obtained in terms of measurable quantities by replacing m_t in the equation of exchange by means of the money demand equation (2):

$$v_t = \beta_1 i_t + (1 - \beta_2) y_t - \varepsilon_t^m \qquad (10)$$

In view of this relationship, money demand shocks and *velocity* shocks are equivalent.

Analogously, equilibrium velocity v_t^* can be defined as a function of the equilibrium nominal interest rate i_t^* and the equilibrium output level y_t^*:

$$v_t^* \equiv \beta_1 i_t^* + (1 - \beta_2) y_t^* \qquad (11)$$

The equilibrium nominal interest rate is given by a Fisher-type identity $i_t^* \equiv r_t^* + (\Delta p)^*_{t+1}$ where $(\Delta p)^*_{t+1}$ denotes the future equilibrium *steady state* inflation rate.

Replacing the liquidity gap in equation (9) by means of equation (10) and (11), the price gap is given in reduced form by

$$p_t^* - p_t = -\beta_1 (i_t - i_t^*) + \beta_2 (y_t - y_t^*) + \varepsilon_t^m \qquad (12)$$

If the price gap in equation (4) is replaced in turn by equation (12), it is evident that the traditional Phillips curve is nested in the inflation equation. While the traditional Phillips curve traces price changes only back to existing output gaps, the liquidity gap within the P-Star model takes into account disequilibria in money holdings in addition to the output gap. These monetary disequilibria take effect on the *current* as well as on the *future* price level. The development of money holdings will become important for the transmission of monetary impulses only if these disequilibria are taken into account.

1.3 The monetary policy rule

The monetary policy rule determines the nominal interest rate and thus endogenises monetary policy. In general, a monetary policy rule may be specified as a feedback rule, according to which monetary policy reacts to deviations of a selected nominal target variable T from a given target value $T^\#$ by appropriately setting the nominal interest rate i given the equilibrium interest rate i^*.

Based on the work of Phillips (1954, 1957), this paper considers a general class of *simple* feedback rules:[10]

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[10] Instead of simple feedback rules, *optimal* feedback rules could be derived given the intertemporal loss function of a monetary authority. However, in the presence of forward-looking expectations the problem of time inconsistency of

$$i_t - i_t^* = \theta_P(T_t - T_t^\#) + \theta_I \sum_{\tau=1}^{t}(T_\tau - T_\tau^\#) + \theta_D \Delta(T_t - T_t^\#) \tag{13}$$

or after taking first differences and simple transformations:

$$\Delta i_t = \Delta i_t^* + (\theta_P + \theta_I + \theta_D)(T_t - T_t^\#) - (\theta_P + 2\theta_D)(T_{t-1} - T_{t-1}^\#) + \theta_D(T_{t-2} - T_{t-2}^\#) \tag{14}$$

(see Blake and Westaway (1994, 1996a)).

While the monetary policy rule is easily implemented using representation (14), the equivalent representation (13) offers an intuitive interpretation of the operation of this general class of policy rules. Representation (13) shows that the deviation of the current interest rate from the equilibrium interest rate depends on three components: the *proportional* (*P*–) component θ_P $(T_t - T_t^\#)$ measures the feedback of the nominal interest rate on the *current* disequilibrium of the target variable, the *integral* (*I*–) component θ_I Σ_τ $(T_\tau - T_\tau^\#)$ the feedback on the *cumulated* disequilibria, and the *differential* (*D*–) component θ_D Δ $(T_t - T_t^\#)$ the feedback on the *change* in the disequilibrium.[11]

In view of the report of the European Monetary Institute on the alternative monetary policy strategies (inflation targeting and monetary targeting), policy rules are specified to control the inflation rate $(\Delta p)_t^\#$ or, alternatively, to control the growth rate of the money stock $(\Delta m)_t^\#$. As an extension of pure inflation targeting and pure monetary targeting, the deviation of current output from a target value $y_t^\#$ is additionally included in the specification of the policy rule. The target value $T^\#$ is then defined

(a) in the case of *pure* or *extended* inflation targeting as:

$T^\# \equiv (\Delta p)^\#$ or $T^\# \equiv ((\Delta p)^\#, y^\#)'$

(b) in the case of *pure* or *extended* monetary targeting as:

$T^\# \equiv (\Delta m)^\#$ or $T^\# \equiv ((\Delta m)^\#, y^\#)'$

The target variable T is defined accordingly. The parameters θ_P, θ_I and θ_D are scalars or vectors of dimension (1×2).

Due to the model's homogeneity in prices, the monetary policy rules guarantee that, given appropriate values of the parameters θ_P, θ_I and θ_D, any inflation rate $(\Delta p)^\#$ or money growth rate $(\Delta m)^\#$ is controllable. For their part, the target inflation rate and the money growth rate determine the equilibrium inflation rate $(\Delta p)^*$.[12] The policy rules impose restrictions on the time path of the inflation rate, but not on the time path of the price level. Nevertheless, the latter can be obtained recursively from the sequence of computed inflation rates given a starting value for the price level.

While the strategy of inflation targeting is aimed directly at the ultimate objective of monetary policy, the strategy of monetary targeting is directed towards controlling the money growth rate, which is an intermediate objective of monetary policy.[13] The target value of the money growth rate is given according to the equation of exchange (8) by:

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optimal monetary policy would be raised. This issue is not under investigation in the present paper. With regard to problems of time inconsistency, see, for example, Blake and Westaway (1994, 1996b).

[11] Specifically, the differential component increases the smoothness of the adjustment path of the target variable to its target level. See Phillips (1957) and Salmon (1982).

[12] Blake and Westaway (1994) show that, given the existence of steady state inflation, the integral component, but not the proportional and differential components, is necessary to control inflation. Taking into account the integral component, however, excludes any *base-drift* in monetary policy.

[13] See Haldane (1995) and Leiderman and Svensson (1995) for a discussion of the concept of inflation targeting and Deutsche Bundesbank (1995) for an exposition of monetary targeting.

$(\Delta m)_t^{\#} = (\Delta p)_t^{\#} + \Delta y_t^* - \Delta v_t^*$

or, if equation (11) is taken into account, by:

$$(\Delta m)_t^{\#} = (\Delta p)_t^{\#} - \beta_1 \Delta i_t^* + \beta_2 \Delta y_t^* \qquad (15)$$

It can be shown within the present model that monetary targeting is a direct generalisation of inflation targeting. From the money demand equation (2) in first differences and the derived target for the monetary growth rate (15), the following relations are immediately obtained using equations (10), (11) and (9):

$$\begin{aligned}\Delta m_t - (\Delta m)_t^{\#} &= (\Delta p_t - (\Delta p)_t^{\#}) - \beta_1(\Delta i_t - \Delta i_t^*) + \beta_2(\Delta y_t - \Delta y_t^*) + \Delta\varepsilon_t^m \\ &= (\Delta p_t - (\Delta p)_t^{\#}) + \Delta(y_t - y_t^*) + \Delta(v_t^* - v_t) \\ &= (\Delta p_t - (\Delta p)_t^{\#}) + \Delta(p_t^* - p_t)\end{aligned}$$

Thus, monetary targeting not only reacts implicitly to a failure to achieve the current target inflation rate. It also responds to changes in the output gap and the liquidity gap, i.e. to changes in the price gap, which is the determinant of current as well as *future* inflation.[14] Hence, monetary targeting is already directed towards future inflationary pressures in a *forward-looking* manner.[15]

2. Model simulation

To illustrate the operation of inflation and monetary targeting within the P-Star model, two scenarios are investigated by means of impulse response analysis:

Scenario 1: A transitory aggregate demand shock.
Scenario 2: A transitory money demand or velocity shock.

The *impulses* of the dynamic system defined by the model are realised in transition from period $t = 0$ to period $t = 1$. The size of these impulses is equal to a unit of output and real money demand, respectively. The adjustment paths of the endogenous variables towards equilibrium, the *responses*, are reported for the periods $t = 1, 2, \ldots, 15$ in deviation from equilibrium[16, 17]

In the first stage, however, the model has to be parameterised. On the basis of empirical findings as well as considerations of plausibility and stability, the parameter values are calibrated at $\alpha_1 = 0.90$, $\alpha_2 = 0.25$, $\alpha_3 = 0.20$, $\beta_1 = 1.00$, $\beta_2 = 0.40$, $\gamma_1 = 0.90$ and $\gamma_2 = 0.20$. The parameter values of the autoregressive equations describing the transition of aggregate demand and velocity shocks are set equal to $\rho^y = \rho^m = 0.50$. As a result of these parameter values, output as well as money demand (in response to serially correlated velocity shocks) exhibit a relatively high degree of persistence. The values of the money demand elasticities are comparable with findings in empirical analyses of money

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14 It is obvious that monetary targeting is equivalent to controlling the equilibrium price level, i.e. P-Star: $\Delta m_t - (\Delta m)_t^{\#} = \Delta p_t^* - (\Delta p)_t^{\#}$.

15 The *current* inflation rate, as a *non-predetermined* variable, reacts immediately to realised shocks, as expectations are formed in a forward-looking way. Therefore, the strategy of inflation targeting, as operationalised here, implicitly takes account of some elements of the *inflation forecast targeting strategy* proposed by Svensson (1997). Should the occasion arise, this strategy might be analysed within the P-Star model by including in the policy rule the deviation of the expected future inflation rate from the target inflation rate $E_t[\Delta p_{t+j}] - (\Delta p)^{\#}_{t+j}$, where j denotes the expectation horizon.

16 The methods applied for solving and simulating the model are described in the Appendix.

17 See Fillion and Tetlow (1994) and Blake and Westaway (1996a) for a description of running stochastic simulations with models under rational expectations.

demand based on broad monetary aggregates. The parameter values of the price equation guarantee that price disequilibria are removed fairly quickly without strongly oscillating adjustment paths.

The parameter values of the policy rule are uniformly chosen with $\theta_P = 0.50$, $\theta_I = 0.50$ and $\theta_D = 0.00$ for the inflation and the monetary growth target deviation as well as for the output disequilibrium. This assignment is motivated by the well-known *Taylor rule* which feeds back the deviation of the current inflation rate from the inflation target and the deviation of current output from equilibrium output to the deviation of the nominal interest rate from its equilibrium value with a parameter value of 0.50 each (see Taylor (1993)). Thus, the policy rule which underlies extended inflation targeting includes the Taylor rule as a special case. Note, however, that the Taylor rule takes into account only the proportional component of the general policy rule (14). The parameter θ_D of the differential component of the policy rule is uniformly set equal to zero as this component dampens the cyclical component of the adjustment paths which, as will be shown below, are already rather smooth without any further dampening.

Bearing in mind the controllability of any inflation rate and any money growth rate, a target inflation rate of $(\Delta p)^\# = 1.00\%$ and a target monetary growth rate of $(\Delta m)^\# = 1.00\%$ are assumed by way of example. Furthermore, if pure inflation and monetary targeting is extended by an output target, the target value chosen for output $y^\#$ is set equal to equilibrium output $y^* = 0$.[18]

2.1 Impulse responses to a transitory aggregate demand shock

Figure 1 below shows the impulse responses of selected endogenous variables to a positive transitory aggregate demand shock for pure inflation and pure monetary targeting. Because of the assumed steady state inflation of 1.00%, nominal levels are transformed into stationary real quantities by subtracting the price level. Both the equilibrium inflation rate and the equilibrium nominal interest rate are equal to one since the real interest rate is set equal to zero. The price gap, the real exchange rate and output are equal to zero in equilibrium, whereas real money demand is equal to minus one.

In response to the serially correlated demand shock a persistent price gap builds up. This induces an increase in the inflation rate, whose value is determined by the current and expected future price gaps in a forward-looking manner. According to the P-Star approach, the price gaps are composed of an output gap and a liquidity gap. Besides the output gap, the liquidity gap, in turn, depends on the deviation of the nominal interest rate from the equilibrium interest rate. In the case of inflation targeting, this deviation is determined by the deviation of current inflation from target inflation, and in the case of monetary targeting by the deviation of the monetary growth rate from its target value.

The real exchange rate, which is determined by the uncovered real interest parity, immediately appreciates. This appreciation as well as the change in the real interest rate feedback to output.[19] The responses of real money demand are determined by the nominal interest rate and by real output. The adjustments of the endogenous variables towards their (initial) equilibrium values take place with time-lags. These lags reflect the transmission mechanism of the model as well as the fact that the serially correlated demand shock diminishes only gradually.

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18 Problems which result if the target value of output is higher than equilibrium output – in this case the realised equilibrium inflation rate is biased upwards – are discussed by Blake and Westaway (1994).

19 The inflation rate and the real exchange rate are non-predetermined variables which immediately jump and put the dynamic system defined by the model on the saddlepoint stable adjustment path. See the exposition in the Appendix.

Figure 1

Impulse responses to a transitory demand shock

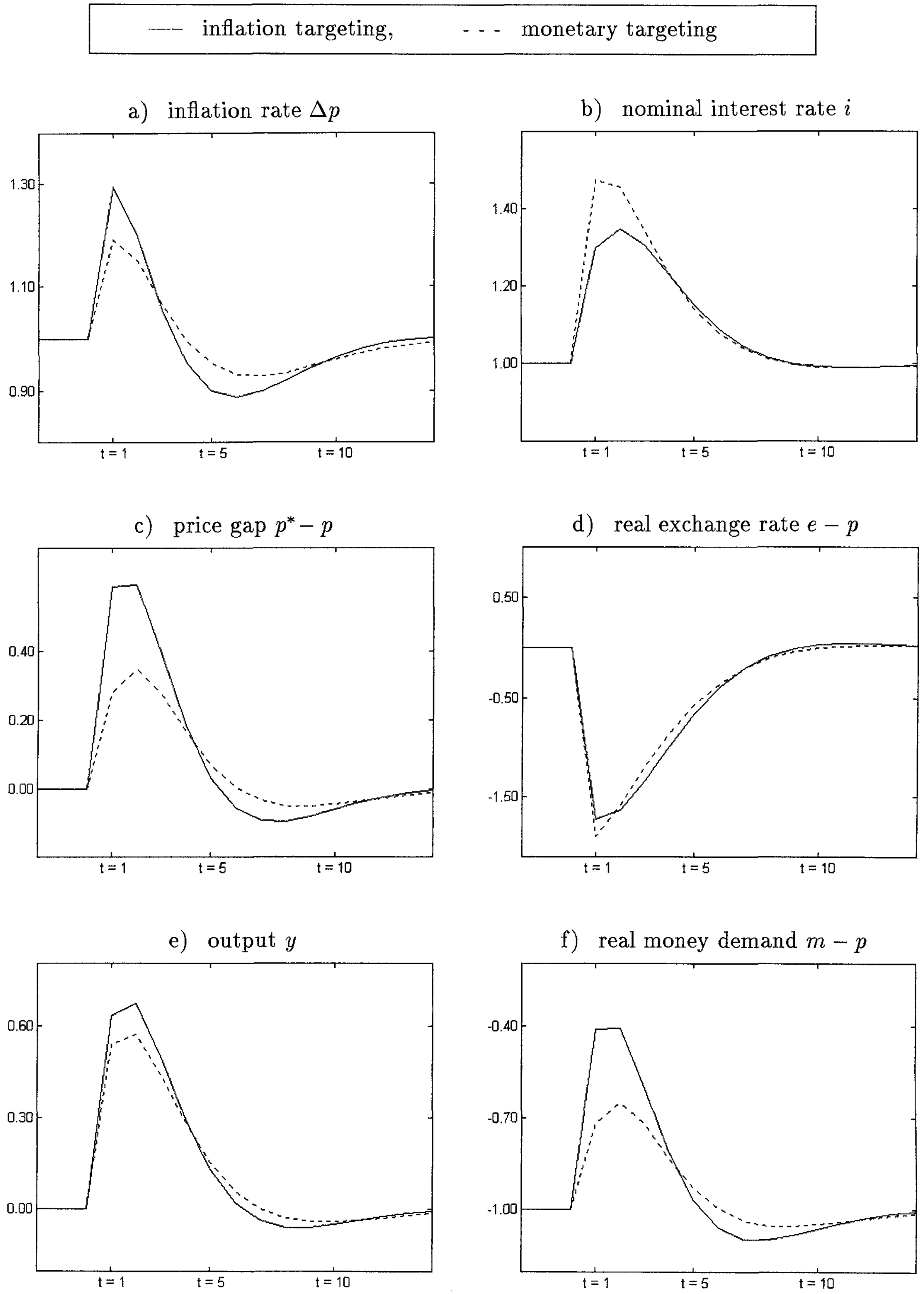


Figure 2

Impulse responses to a transitory demand shock

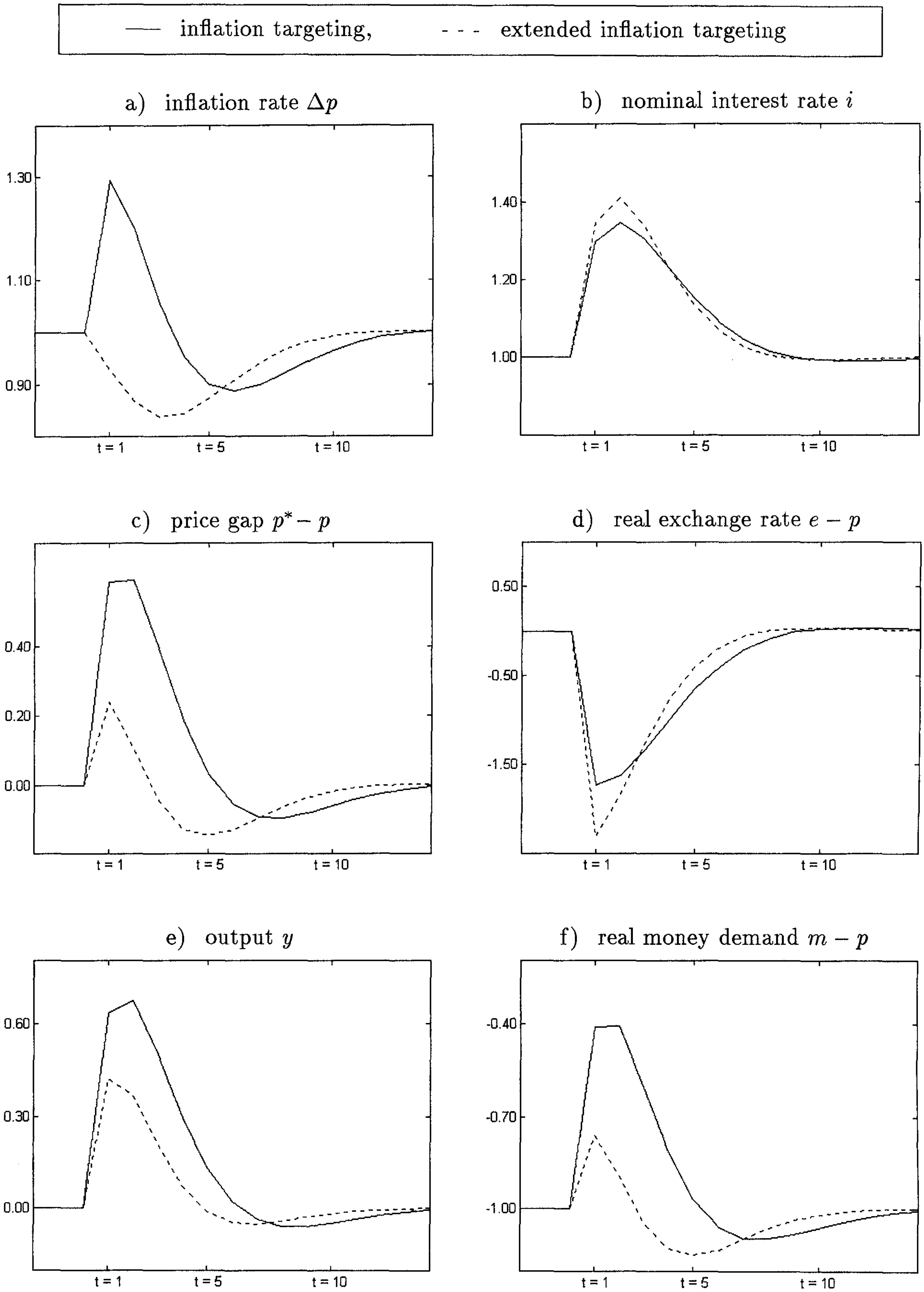


Figure 3

Impulse responses to a transitory demand shock

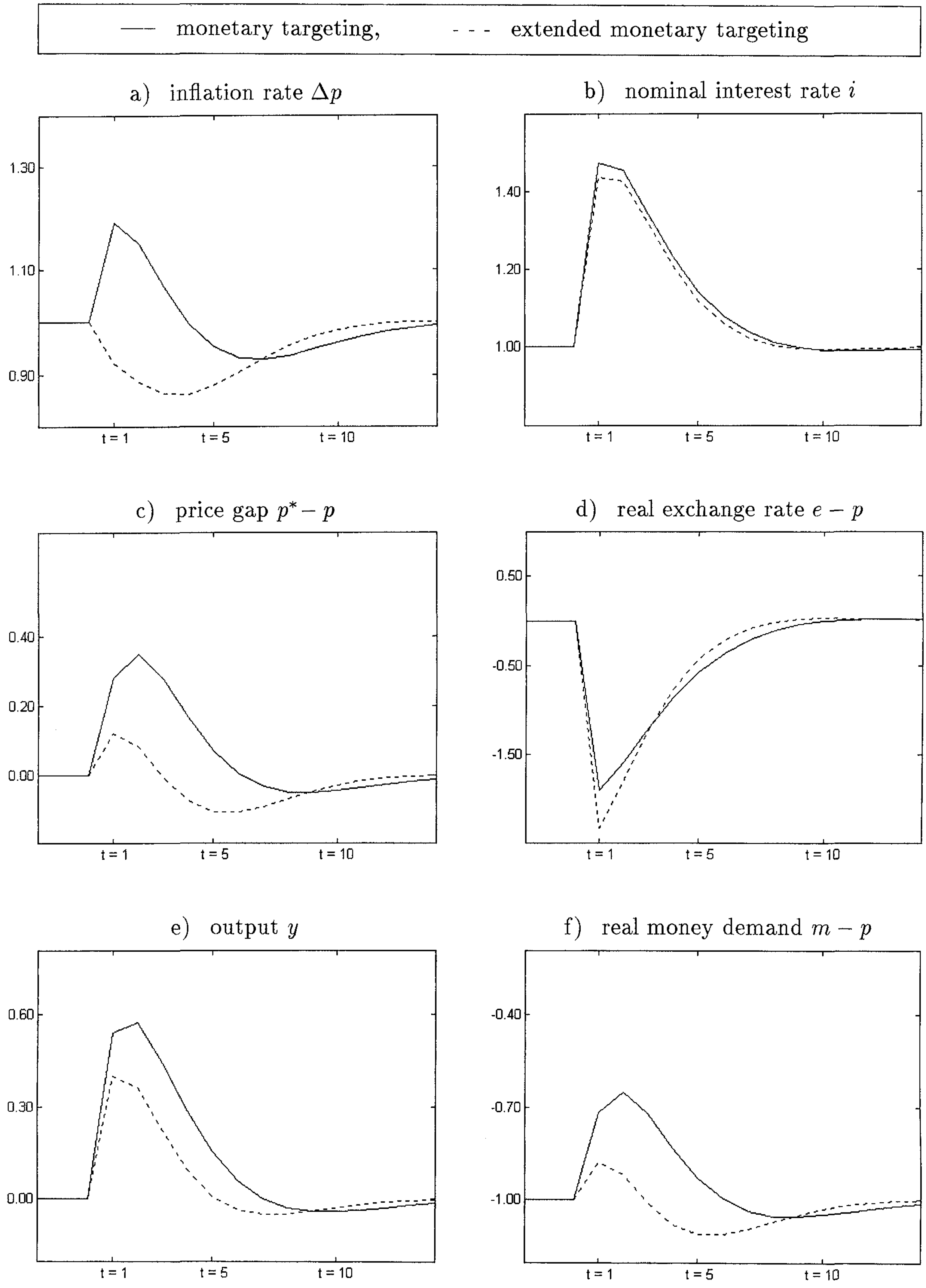


Figure 4

Impulse responses to a transitory velocity shock

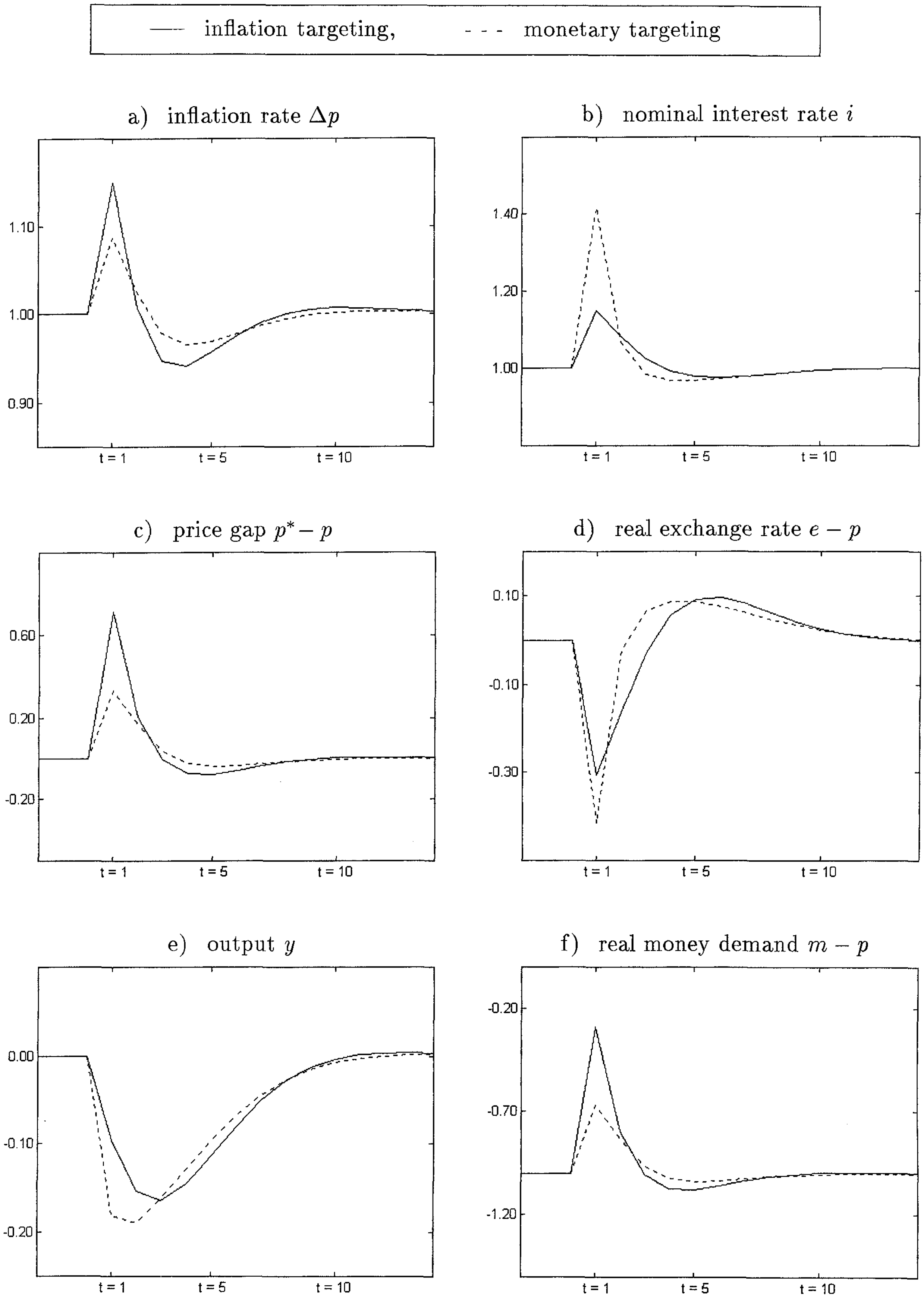


Figure 5

Impulse responses to a transitory velocity shock

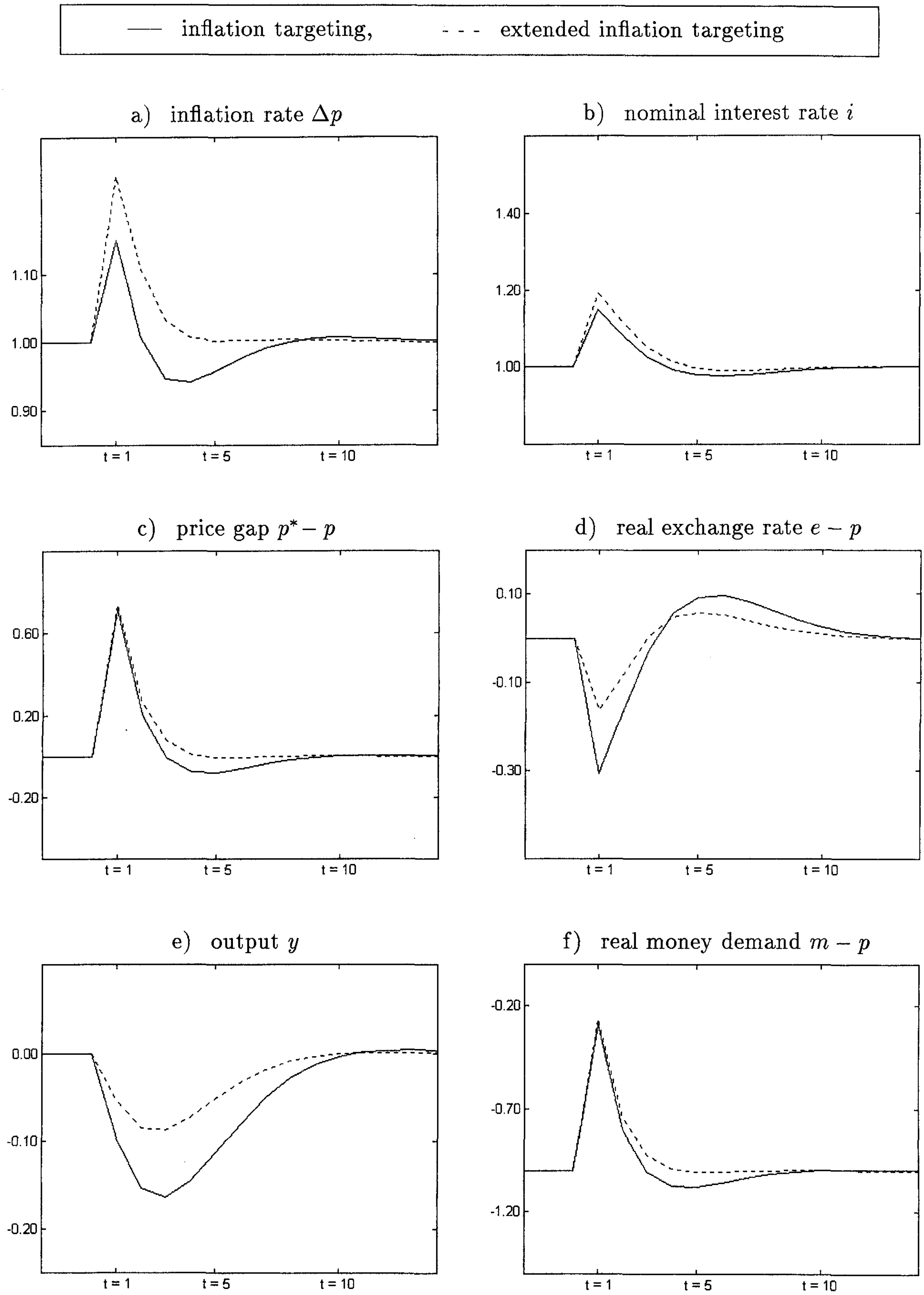
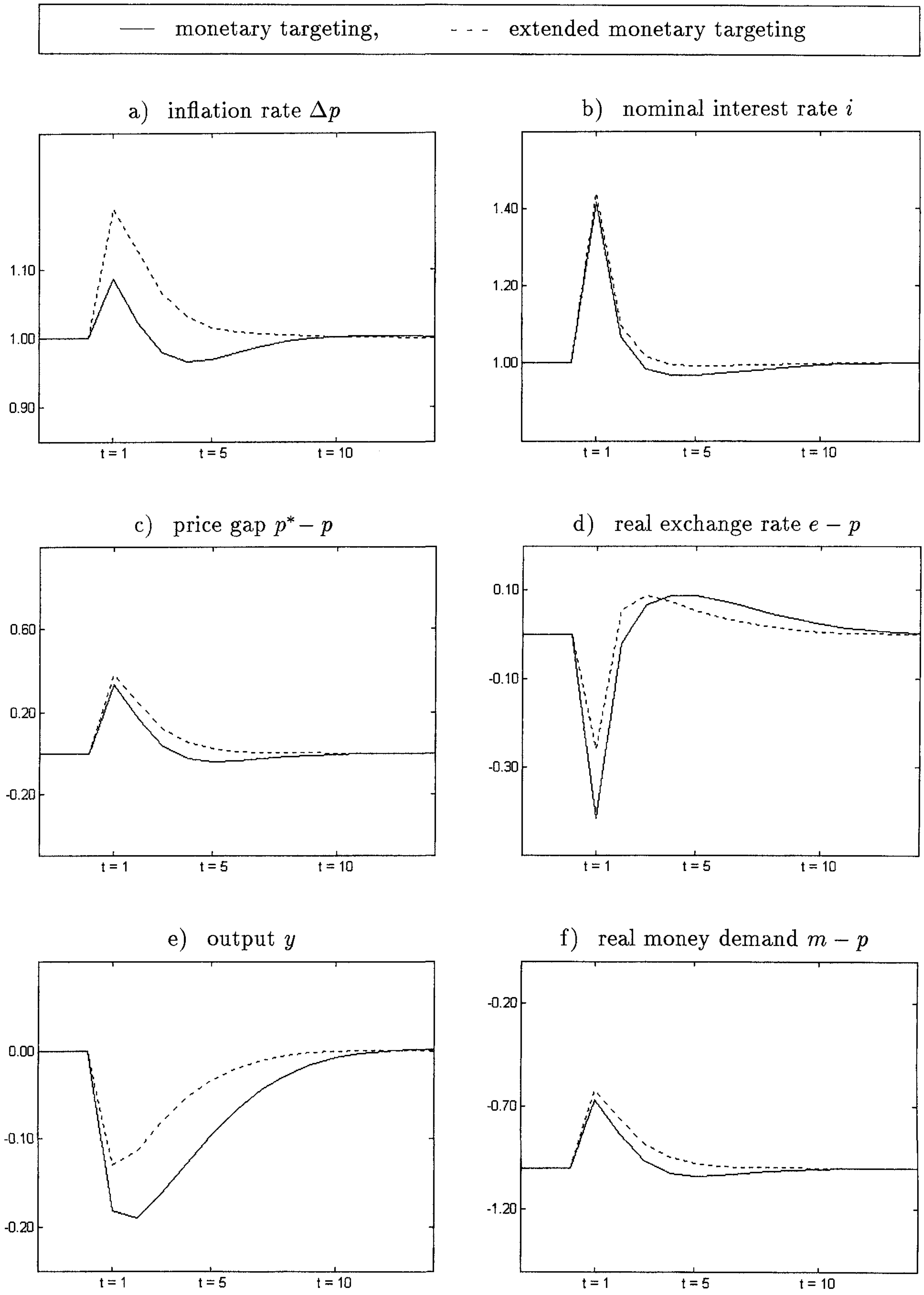


Figure 6

Impulse responses to a transitory velocity shock



If the adjustment paths of the endogenous variables are compared, it is obvious that the inflation rate displays lower volatility in the case of monetary targeting than it does in the case of inflation targeting. This result, however, is accompanied by a stronger response of the nominal interest rate, i.e. the monetary policy instrument. Inflationary pressures are weaker in the case of monetary targeting due to the stronger interest rate response; this reflects the fact that changes in the liquidity and output gap are taken into account in addition to the current inflation disequilibrium as shown in sub-section 1.2. Thus, the immediate increase in inflation is smaller. On the other hand, the stronger interest rate response induces a larger real appreciation.

Figures 2 and 3 show the impulse responses of the endogenous variables to the transitory demand shock for extended inflation and monetary targeting together with the impulse responses in the case of pure inflation and pure monetary targeting.

It is evident that letting the nominal interest rate depend on output disequilibria results in a faster return to equilibrium. This result holds for extended inflation targeting as well as for extended monetary targeting. However, it is accompanied by a transitory *decrease* in the inflation rate. This adverse reaction reflects the extreme assumptions underlying the forward-looking inflation expectations and, in particular, the selected parameter values in the inflation equation which heavily weights the future negative price gaps.

Accordingly, a *negative* demand shock or a business cycle trough would be countered by an interest rate decrease, i.e. an expansionary monetary reaction. This reaction would induce a transitory *increase* in the inflation rate. Hence, in the light of this finding, the extension of the monetary policy rule by output disequilibria should be judged critically as monetary policy is obliged to give priority to price stability.[20]

2.2 Impulse responses to a transitory money demand shock

The operation of inflation and monetary targeting in response to a transitory money demand or velocity shock is shown in Figure 4. The equilibrium values are identical to those of Scenario 1.

Owing to the serially correlated velocity shock, a persistent price gap emerges that induces an increase in the inflation rate. The inflationary impulse is counteracted by monetary policy by increasing the nominal interest rate according to the respective policy rule. Monetary targeting again responds by increasing the interest rate more sharply than would have been the case under inflation targeting and can thereby check inflationary impulses to a greater degree through a larger reduction in the liquidity gap. Analogous to the operation of inflation and monetary targeting in response to an aggregate demand shock, monetary targeting is again characterised by a lower volatility of the inflation rate. Furthermore, the volatility of the nominal interest rate is again higher.

Figures 5 and 6 show the impulse responses in the case of inflation and monetary targeting extended by output disequilibria.

Of course, in comparison with pure inflation and pure monetary targeting the feedback of interest rate changes on the output gap accelerates the reduction of output disequilibria. At the same time, the inflation rate increases more strongly. In other respects, however, the impulse responses do not differ fundamentally.

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20 If the Taylor rule including output disequilibria is compatible with empirical findings this result might be explained by the fact that monetary policy has historically been repeatedly used to stabilise business cycles.

Concluding remarks

In the present paper, the operation of inflation and monetary targeting has been analysed using a model of a small open economy based on the P-Star approach. The monetary policy regimes under investigation are the two alternative monetary policy strategies considered for the European System of Central Banks in Stage III of the European Monetary Union. Model-based simulations show that under a regime of monetary targeting the inflation rate has a lower volatility in response to demand shocks as well as in response to velocity shocks than under a regime of inflation targeting. The lower volatility of the inflation rate, however, is accompanied by a higher volatility of the nominal interest rate and, hence, of the exchange rate determined on the basis of uncovered interest parity. Thus, within the P-Star model, where inflation is a monetary phenomenon in the long run, there is much to be said for monetary targeting aimed at controlling the long-term determinant of inflation, i.e. the money stock.

If the simplifying assumptions which underlie the theoretical analysis are set aside, monetary policy makers are faced with the practical problem of operationalising monetary targeting. On the one hand, given a stable money demand, i.e. velocity is forecastable (a key assumption which underlies the specification of the P-Star model), monetary targeting ensures the controllability of money holdings with a fairly high degree of reliability. At the same time, monetary targeting offers a high degree of transparency to the general public. This transparency results not least from the timely availability of data on the current development of monetary aggregates. On the other hand, the recent instability of the financial sector in many countries renders the realisation of the ultimate goal of price stability by monetary targeting more difficult.

Inflation targeting which directly aims at the ultimate goal of price stability is often motivated by the failure of monetary targeting due to the instability of the financial sector. However, problems operationalising inflation targeting result from measuring inflation which is feasible only with a time-lag and which suffers from non-uniqueness. Furthermore, long and variable time-lags have to be taken into account when using monetary policy instruments to control inflation directly. Bearing that in mind, it would be advantageous to base inflation targeting on the *expected* future inflation rate instead of the *current* inflation rate, i.e. to follow a strategy of *inflation forecast targeting* which is proposed by Svensson (1997) in particular. However, inflation forecast targeting entails the problem of forecasting inflation with sufficient accuracy. As yet, this problem has not been tackled successfully.

Regarding the extension of pure inflation and pure monetary targeting by output disequilibria, it has to be pointed out that data on current output are only available with a time-lag and that the development of equilibrium output is uncertain. Furthermore, taking an output target into account could threaten the independence of monetary policy, whose main priority should, after all, be price stability.

Given the fact that the analysis in the present paper is confined to a stylised calibrated model, it has to be stressed that the analysis should be placed on a stronger empirical footing if it is to contribute to the discussion on the design of monetary policy beyond the theoretical findings documented here. Only a model which is firmly based on empirical grounds will provide a reliable framework for contrasting the operation of monetary and inflation targeting.

In particular, the parameterisation of the inflation equation and the monetary policy rule, which essentially determine the dynamic properties of the monetary transmission mechanism, needs further investigation. Against this background, the model under investigation should be estimated or at least calibrated taking a statistical criterion as a basis. Subsequently, the model could be evaluated using stochastic simulations to ascertain how far it matches empirical regularities measured in the data.

Appendix: solving and simulating the model

The solution of the model described in Section 1 is obtained using the method suggested by Blanchard and Kahn (1980).[21] Initially, the structural equations (1) – (5), (6), (12) and the policy rule (14) are written in *state space form*:

$$\begin{pmatrix} x_{1,t+1} \\ E_t\left[x_{2,t+1}\right] \end{pmatrix} = A\begin{pmatrix} x_{1,t} \\ x_{2,t} \end{pmatrix} + B\eta_{t+1} \tag{16}$$

with the *state vector* $x_t = (x'_{1,t}, x'_{2,t})'$ and the *transition matrix*

$$A = \begin{pmatrix} A_{11} & A_{12} \\ A_{21} & A_{22} \end{pmatrix}$$

which is partitioned according to the dimension of the state vectors $x_{1,t}$ and $x_{2,t}$, where

(a) in the case of extended inflation targeting:

$$x_{1,t} = \left((\Delta p)^{\#}_t, (\Delta p)^{\#}_{t-1}, (\Delta p)^{\#}_{t-2}, i_{t-1}, y_{t-1}, y_{t-2}, \varepsilon^y_t, m_{t-1} - p_{t-1}, \varepsilon^m_t, r_{t-1}, p^*_{t-1} - p_{t-1}, \Delta p_{t-1}, \Delta p_{t-2}\right)'$$

$$x_{2,t} = \left(e_t - p_t, \Delta p_t\right)'$$

(b) in the case of extended monetary targeting:

$$x_{1,t} = \left((\Delta m)^{\#}_t, (\Delta m)^{\#}_{t-1}, (\Delta m)^{\#}_{t-2}, i_{t-1}, y_{t-1}, y_{t-2}, \varepsilon^y_t, m_{t-1} - p_{t-1}, m_{t-2} - p_{t-2}, \varepsilon^m_t, r_{t-1}, \Delta m_{t-1}, \Delta m_{t-2}, \right.$$
$$\left. p^*_{t-1} - p_{t-1}, \Delta p_{t-1}, \Delta p_{t-2}\right)'$$

$$x_{2,t} = \left(e_t - p_t, \Delta p_t\right)'$$

The *input matrix* $B = (B'_1, B'_2)'$ is partitioned according to the dimension of the state vectors $x_{1,t}$ and $x_{2,t}$ taking into account the dimension of the innovation $\eta_{t+1} = (\eta^y_{t+1}, \eta^m_{t+1})'$.

The vector $x_{1,t}$ contains the *predetermined* state variables of period t, the vector $x_{2,t}$ the *non-predetermined* state variables of period t. Non-predetermined are those variables whose realisations in the future period $t+1$ are subject to forward-looking expectations based on the information set Ω_t available in period t.[22] Thus, within the model under consideration, the (real) exchange rate and the inflation rate are non-predetermined irrespective of the monetary regime.

When writing the model in state space form, it has to be borne in mind that the nominal levels are trending by reason of the assumed steady state inflation. Therefore, as the method of Blanchard and Kahn presupposes the existence of a *stationary* equilibrium, these variables have to be transformed into stationary quantities by subtracting the price level, as is already shown by the definition of the state vectors under (a) and (b).[23] A necessary condition for this transformation is the homogeneity of the model in the price level.

The solution of the dynamic equation system (16) is *saddlepoint stable*, i.e. uniquely stable, if the number of the eigenvalues of the transition matrix A which lie outside the (complex) unit circle equals the number of the non-predetermined state variables and the number of the eigenvalues of the transition matrix A which lie inside the unit circle equals the number of the predetermined state variables (see Blanchard and Kahn (1980), Proposition 1).

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[21] See also the exposition in Buiter (1984, 1986).

[22] See Buiter (1982) for his amendment to the definition given by Blanchard and Kahn (1980).

[23] See Buiter and Miller (1982).

The condition of saddlepoint stability is satisfied for the model irrespective of the monetary regime given the parameter values of Section 2 In view of two non-predetermined variables – the (real) exchange rate and the inflation rate – two (complex conjugate) eigenvalues lie outside the unit circle.

The saddlepoint stable solution of the model has to be determined on the basis of the information set Ω_t available in period t. If the conditional expectation operator $E_t[\cdot]$ is applied to the equation system (16) and account is taken of the fact that $E_t[x_t] = x_t$ and $E_t[\eta_{t+1}] = 0$, the following (deterministic) equation system is obtained:[24]

$$E_t[x_{t+1}] = Ax_t \tag{17}$$

The transition matrix A is transformed into the Jordan canonical form:

$$A = V \Lambda V^{-1} \tag{18}$$

with:

$$\Lambda = \begin{pmatrix} \Lambda_1 & 0 \\ 0 & \Lambda_2 \end{pmatrix}, \quad V = \begin{pmatrix} V_{11} & V_{12} \\ V_{21} & V_{22} \end{pmatrix}, \quad V^{-1} = \begin{pmatrix} W_{11} & W_{12} \\ W_{21} & W_{22} \end{pmatrix}$$

where the matrices Λ, V and V^{-1} are partitioned to conform with the partition of the vector x_t. Assuming non-repeated eigenvalues, Λ is a diagonal matrix with the eigenvalues of the transition matrix A on its main diagonal; the matrix V is a matrix whose column vectors are the right eigenvectors corresponding to the eigenvalues and the matrix V^{-1} is a matrix whose row vectors are the left eigenvectors corresponding to the eigenvalues (see Golub and van Loan (1989), p. 339). The eigenvectors are ordered in such a way that the eigenvalues of the diagonal matrix Λ_1 lie inside the unit circle and the eigenvalues of the diagonal matrix Λ_2 lie outside the unit circle.

If the matrix of the left eigenvectors V^{-1} is multiplied from the left, the equation system (17) can be transformed into a system in the canonical variables $\tilde{x}_t = (\tilde{x}'_{1,t}, \tilde{x}'_{2,t})'$:

$$E_t\left[\tilde{x}_{t+1}\right] = \Lambda \tilde{x}_t \tag{19}$$

with $\tilde{x}_t \equiv V^{-1} x_t$.

Owing to the diagonal structure of matrix Λ, the transformed system (19) is decoupled. Hence, the subsystems:

$$E_t\left[\tilde{x}_{i,t+1}\right] = \Lambda_i \tilde{x}_{i,t} \quad i = 1,2$$

can be solved independently from each other.

As the eigenvalues on the main diagonal of Λ_2 lie outside the unit circle, the stable solution of the corresponding subsystem is to be determined by forward substitution of:

$$\tilde{x}_{2,t} = \Lambda_2^{-1} E_t\left[\tilde{x}_{2,t+1}\right] \tag{20}$$

After repeated substitution and application of the law of iterated expectations the solution is given by:

$$\tilde{x}_{2,t} = \lim_{\tau \to \infty} \Lambda_2^{-\tau-1} E_t\left[\tilde{x}_{2,t+\tau+1}\right]$$

It immediately follows that $\lim_{\tau \to \infty} \Lambda_2^{-\tau-1} E_t\left[\tilde{x}_{2,t+\tau+1}\right] = 0$ and thus:

[24] The solution for the vector of the non-predetermined variables $x_{2,t}$ is restricted to the class of linear functions of the vector of the predetermined variables $x_{1,t} \in \Omega_t$. Thus, the vector $x_{2,t}$ is implicitly an element of the information set Ω_t too.

$$\tilde{x}_{2,t} = 0 \tag{21}$$

Starting from (21) and taking into account the relationship $\tilde{x}_{2,t} = W_{21}x_{1,t} + W_{22}x_{2,t}$ the following result is obtained:

$$x_{2,t} = -W_{22}^{-1}W_{21}x_{1,t} \tag{22}$$

Thus, the vector of non-predetermined variables $x_{2,t}$ is given by a time-invariant linear function of the vector of predetermined variables $x_{1,t}$ depending on the left eigenvectors which correspond to the eigenvectors of A lying outside the unit circle.

If $x_{2,t}$ in system (16) is replaced by means of equation (22), the transition equation of the state vector $x_{1,t}$ is given by:

$$x_{1,t+1} = (A_{11} - A_{12}W_{22}^{-1}W_{21})x_{1,t} + B_1\eta_{t+1} \tag{23}$$

Reconsidering the decomposition of the partitioned transition matrix A according to (18), it follows from the formulae for inverting partitioned matrices (see Graybill (1983), p. 184) that:

$$\begin{aligned} A_{11} - A_{12}W_{22}^{-1}W_{21} &= V_{11}\Lambda_1(W_{11} - W_{12}W_{22}^{-1}W_{21}) \\ &= V_{11}\Lambda_1 V_{11}^{-1} \end{aligned}$$

Hence, the transition equation (23) can be written equivalently as:

$$x_{1,t+1} = V_{11}\Lambda_1 V_{11}^{-1}x_{1,t} + B_1\eta_{t+1}$$

Obviously, this transition equation is stable as the eigenvalues on the main diagonal of Λ_1 lie inside the unit circle.

By renewed application of the formulae for inverting partitioned matrices it can be shown that the identity:

$$-W_{22}^{-1}W_{21} = V_{21}V_{11}^{-1}$$

holds. Thus, the linear function (22) can be alternatively obtained using the right eigenvectors which correspond to the eigenvalues of A lying inside the unit circle.

If the preceding results are combined, the solution of the state space model (16) is:

$$x_{1,t} = Mx_{1,t-1} + B_1\eta_t \tag{24}$$

$$x_{2,t} = Nx_{1,t} \tag{25}$$

with $M \equiv V_{11}\Lambda_1 V_{11}^{-1}$ and $N \equiv V_{21}V_{11}^{-1}$, where the transition equation of the predetermined variables given by (24) is shifted back in time one period.

If the solution formulae (24), and (25) are employed and the predetermined variables x_1 are given appropriate starting values $x_{1,0}$, model (16) can be easily simulated for $t = 1, 2, \ldots$ given a sequence of innovations η_t, $t = 1, 2, \ldots$. Here, the non-predetermined variables x_2 *jump* in each period $t = 1, 2, \ldots$ to reach a level $x_{2,t}$ that puts the vector of predetermined and non-predetermined variables $x_t = (x'_{1,t}, x'_{2,t})'$ on the saddlepoint stable adjustment path.

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Comments on: "Inflation versus monetary targeting in a P-Star model with rational expectations" by Günter Coenen

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by Peter J. A. van Els[1]

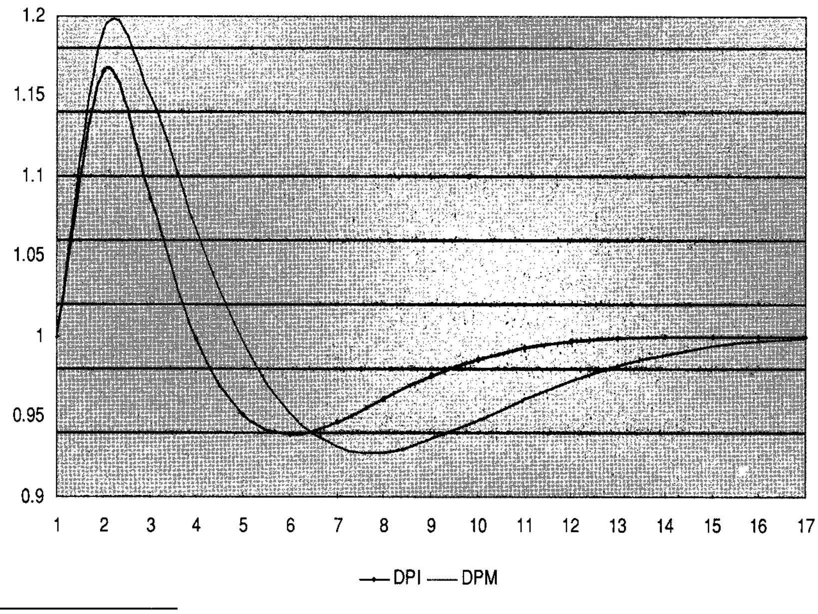
Günter Coenen has written an insightful paper on the issue of inflation versus monetary targeting. The main contribution of Coenen's paper is that, in exploring the pros and cons of both monetary policy strategies, he uses a model which takes into account an explicit role for money in controlling inflation, the so-called P-Star model. Within this framework a strategy aimed at controlling money growth is not inefficient *a priori*, and hence may be compared with alternative monetary policy strategies. The topic of inflation versus monetary targeting is an important one, as the ESCB considers these two strategies to be the only possible strategies to be followed in stage III of EMU.

My comments are fourfold.

First, although the approach followed by Coenen could be seen as a fruitful and promising one in principle, his main points that "... *under a regime of monetary targeting the inflation rate has a lower volatility [...] than under a regime of inflation targeting...*" and "*...within the P-Star model [...] there is much to be said for monetary targeting ...*" are unwarranted as general conclusions. These results depend crucially on the parameter values of the policy feedback rules. In principle there is no reason why these parameter values should be identical under both strategies, as Coenen assumes in his paper. Starting from a framework in which monetary policy explicitly aims at price stability, and given the restrictions, economic conditions, the monetary policy strategy, the parameters of reduced-form feedback rules will, in general, depend on the strategy chosen. In other words, monetary authorities will take into account that inflationary shocks require a different policy response under different policy regimes to achieve the same goal. An illustration of this is given in the graph below, which shows the responses of inflation to a positive demand shock under inflation

Responses of inflation to a positive demand shock

Percentage points



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targeting with θ_l =1.5 (DPI) and monetary targeting with θ_l =0.5 (DPM), respectively. In this specific case inflation targeting results in a lower rise of inflation than monetary targeting. Hence, the conclusion seems to be that within the framework of Coenen's model the same inflation performance can be achieved with both monetary policy strategies.

Second, the issue of controllability of the money stock is addressed only very rudimentary. Given the fact that money demand is inversely related to the short-term interest rate, the money stock *m* must reflect some form of narrow money, not M3. However, empirical evidence for Germany suggests that the price gap in terms of M3 is a more reliable predictor of inflation than M1.[2] Introducing M3 would change the analysis to the extent that the demand for M3 and the short-term interest rate are typically positively related. This would call for incorporating the short-term as well as the long-term interest rate in the money demand equation.[3] Moreover, the long-term rate must be made endogenous through some form of term structure relationship. This would give the model a stronger empirical basis. However, controllability of money is no longer guaranteed *a priori*.[4]

Third, it is not clear why the analysis is placed within the small open economy framework. Judged by trade to GDP ratios, the EMU area is neither a very open economy; nor a small economy. Moreover, whereas the model allows the real exchange rate to affect output directly, prices are not influenced by exchange rate changes, which is inconsistent with the small open economy assumption. Some own experiments with Coenen's model show that if inflation is affected by exchange rate changes directly, the impact response of inflation to a positive demand shock could be negative rather than positive. All in all, the question rises what to make of the small open economy assumption and whether it would not be better to conduct the analysis within a two-country framework.

Finally, the question rises whether it will be possible to choose between inflation and monetary targeting on the basis of stylised model exercises. I fully agree with the author that a reliable model framework should be firmly based on empirical foundations. Some evidence suggests that in practice the differences between monetary and inflation targeting should not be exaggerated. For instance, past changes in official German interest rates have been found to be related not only to deviations from announced money growth targets but even more significantly to inflation performance and output gap movements.[5] However, one should bear in mind that a number of very important issues at stake in the choice between alternative strategies will be difficult to capture in models. Here one could think of issues such as protection against political pressures, the process of gaining and establishing credibility, accountability, and transparency.

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2 See for instance K.-H. Tödter and H.-E. Reimers (1994): "P-Star as a link between money and prices in Germany", *Weltwirtschaftliches Archiv* 130, pp. 273-89, and J. M. Groeneveld (1997): *Inflation Dynamics and Monetary Strategies*, PhD Dissertation University of Maastricht, Thesis Publishers Amsterdam.

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Interpreting a Monetary Conditions Index in economic policy

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Introduction

The main purpose of this paper is to review and interpret the use of a Monetary Conditions Index (or MCI) by central banks in the conduct of monetary policy. Numerous central banks, governmental organizations, and businesses now calculate an MCI as an indicator of the stance of monetary policy. Two central banks, those for Canada and New Zealand, use their MCIs as operational targets.

This paper describes and defines the concept of an MCI, summarizes how central banks implement MCIs in practice, reviews some of the operational and conceptual issues involved, and evaluates the sensitivity of MCIs to an inherent source of uncertainty in their calculation. Empirically, this uncertainty typically results in MCIs that are uninformative as indicators of monetary conditions, so some possible alternatives are briefly considered.

1. A Monetary Conditions Index in practice

Several central banks calculate a Monetary Conditions Index for use in monetary policy. Empirically, an MCI is a weighted average of changes in an interest rate and an exchange rate relative to their values in a base period. The weights on the interest rate and exchange rate reflect the estimated relative effects of those variables on aggregate demand over some period, often approximately two years. MCIs are currently used as indicators of monetary conditions and as operational short-run targets for monetary policy.

A Monetary Conditions Index has several attractive features. Its motivation is simple: exchange rates influence aggregate demand, especially in small open economies. Thus, focusing on exchange rates as well as interest rates may be important in understanding an economy's behavior, and so in policymaking. Also, an MCI is easy to calculate. For central banks, an MCI is an intuitively

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appealing operational target for monetary policy. It generalizes interest-rate targeting to include effects of exchange rates on an open economy, and it serves as a model-based policy guide between formal model forecasts. For institutions other than central banks, an MCI as an index per se may capture both domestic and foreign influences on the general monetary conditions of a country.

MCIs have gained widespread use. The central banks of Canada, New Zealand, Norway, and Sweden each have published an MCI and, to varying degrees, use their respective indexes in the conduct of monetary policy. Additionally, the International Monetary Fund (IMF) and the Organisation for Economic Co-operation and Development (OECD) calculate MCIs for evaluating the monetary policies of many countries; and firms such as Deutsche Bank, Goldman Sachs, JP Morgan, and Merrill Lynch publish MCIs to ascertain the general monetary environment in various countries.

An MCI assumes an underlying model relating economic activity and inflation to the variables in the MCI, with the weights in the MCI reflecting the effects of the interest rate and exchange rate on aggregate demand. Being model-based, those variables' effects are estimated, and the corresponding coefficients have an associated uncertainty from estimation. This paper shows that, empirically, this uncertainty typically renders MCIs uninformative for their ostensible purposes.

Sections 2 and 3 provide a foundation for understanding MCIs in practice, and hence for understanding how estimation uncertainty impinges on their use. Section 2 describes and defines the Bank of Canada's Monetary Conditions Index, summarizes how the Bank utilizes its MCI in conducting monetary policy, reviews some operational considerations, and documents MCI usage by other institutions and for other countries. Section 3 analyzes two facets in the design of an MCI: the choice of weights and variables, and the assumptions of the underlying empirical model. Section 4 presents confidence intervals of estimated relative MCI weights derived from models for Canada, New Zealand, Norway, Sweden, and the United States. In light of the (often) extreme uncertainty present in calculating MCIs, Section 4 then considers some possible alternatives. While intuitively appealing, an MCI appears fraught with difficulties as an indicator of monetary stance and as an operational target for monetary policy.

A previous paper, Eika, Ericsson, and Nymoen (1996a), derives analytical and empirical properties of MCIs in an attempt to ascertain their usefulness in monetary policy. The current paper complements Eika, Ericsson, and Nymoen (1996a) by focusing on the practical implementation of an MCI and the degree to which implementation is affected by uncertainty in the estimated weights.

2. Construction and use of MCIs

The concept of a Monetary Conditions Index was developed at the Bank of Canada and has been used there more extensively than elsewhere, so this section begins by describing the Bank's MCI (Section 2.1), its implementation in practice (Section 2.2), and operational considerations (Section 2.3). Section 2.4 considers MCIs used by other institutions and for other countries. The discussion in the first three subsections relies heavily on Duguay and Poloz (1994), Poloz, Rose, and Tetlow (1994), and Longworth and Freedman (1995) for the role of the Bank's quarterly model in monetary policy; on the Bank of Canada (1994, 1995), Barker (1996), and Zelmer (1996) for details on the MCI itself; on Freedman (1994) for the justification of an MCI in monetary policy; and especially on Freedman (1995) and Thiessen (1995) for overviews encompassing all of these issues.

2.1 Construction of the Bank of Canada's MCI

For the last several years, the Bank of Canada has used an MCI as an operational target in setting monetary policy. This subsection defines the construction of the MCI, briefly describes its empirical underpinnings, and interprets the generated index.

The Bank's MCI is a weighted sum of changes in the nominal Canadian 90-day commercial paper interest rate (R) and a nominal G-10 bilateral trade-weighted exchange rate index

(E), where both variables are relative to values in a base period. The weights on the interest rate and exchange rate reflect their estimated relative effects on Canadian output. The Bank of Canada uses weights of 3 to 1, interest rate to exchange rate. That is, a one percentage point increase in the interest rate induces three times the change in the Bank's MCI as would a 1% appreciation of the Canadian dollar. Algebraically, it is convenient to write the MCI as:

$$MCI_t = \theta_R (R_t - R_0) + \theta_e (e_t - e_0), \tag{1}$$

where t is a time index, $t = 0$ is the base period, θ_R and θ_e are the respective weights on the interest rate and the exchange rate, and variables in lower case denote logarithms. Thus, the calculated MCI depends upon the weights θ_R and θ_e, the measures of the exchange rate and the interest rate, and the base period. Usually, the exchange rate in (1) is in logarithms or in percent deviations from its baseline value, whereas the interest rate is in levels. Below, logarithms of the exchange rate generally are used, and the choice makes little difference for the countries and sample periods involved.

The relative weight of 3 is derived from a range of econometric evidence on the determinants of aggregate demand. As discussed in Freedman (1994, pp. 469–70 and footnote 27), Duguay's (1994) results are typical of that evidence, so we focus on a representative regression from that paper, Duguay (1994, p. 50, Table 1, column 7):

$$\begin{aligned} \Delta y_t = \ & \underset{(0.13)}{+0.13} + \underset{(0.11)}{0.52}\Delta y_t^* + \underset{(0.11)}{0.45}\Delta y_{t-1}^* - \underset{(0.22)}{0.40}[\Delta_8 RR_t / 8] - \underset{(0.12)}{0.15}[\Delta_{12} q_t / 12] \end{aligned} \tag{2}$$

$$T = 44[1980(1) - 1990(4)] \qquad \bar{R}^2 = 0.64 \qquad \hat{\sigma} = 0.62\% \qquad dw = 1.96.$$

The series are all quarterly and include real Canadian GDP (Y) and real US GDP (Y^*); and Δ is the first difference operator.[1] The real interest rate *(RR)* is constructed as the nominal 90-day commercial paper interest rate (R) minus the one-quarter lag in the annual rate of change of the Canadian GDP deflator (P). That is, $RR_t = R_t - \Delta_4 p_{t-1}$. The real exchange rate (Q) is the product of the nominal bilateral US-Canadadian exchange rate (E, in US dollars per Canadian dollar) and the ratio of the Canadian GDP deflator to the US GDP deflator (P^*): i.e., $Q = E \bullet (P/P^*)$. Thus, an increase in Q represents an appreciation of the Canadian real exchange rate. The symbols T, $\bar{R}^2$, $\hat{\sigma}$, and dw denote the sample size of the estimation period, the adjusted squared multiple correlation coefficient, the estimated equation standard error, and the Durbin-Watson statistic respectively. The coefficients are estimated by least squares, and estimated standard errors are in parentheses.[2]

In (2), the ratio of the coefficients on the interest rate and the exchange rate is (–0.40)/(–0.15) or 2.67, which is virtually the relative weight of 3 used by the Bank of Canada. While the relative weight is based on estimated relations with real interest rates and exchange rates, the Bank applies the weight to an index with the corresponding nominal variables. Switching from real to nominal variables is convenient operationally, and it has been defended by the short horizon for MCI-based monetary policy and the near constancy of inflation and relative prices over that horizon.

Figure 1 plots the Bank's Monetary Conditions Index. A decline in the interest rate increases aggregate demand and lowers the MCI, as does a depreciation of the Canadian dollar, so a

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1 The difference operator Δ is defined as $(1\text{-}L)$, where the lag operator L shifts a variable one period into the past. Hence, for x_t (a variable x at time t), $Lx_t = x_{t-1}$ and so $\Delta x_t = x_t - x_{t-1}$. More generally, $\Delta_j^i x_t = (1-L^j)^i x_t$. If i (or j) is undefined, it is taken to be unity.

2 A minor notational and empirical discrepancy exists between (1) and (2), in that E in the former is the G-10 trade-weighted exchange rate whereas E (through Q) in the latter is the bilateral US-Canadian exchange rate. This distinction is maintained below. MCIs for Canada use the G-10 trade-weighted exchange rate, whereas regressions for Canada use the bilateral US-Canadian exchange rate. Choice between the two exchange rates should make only a minor difference: the US-Canadian exchange rate dominates the G-10 trade-weighted exchange rate, with the former receiving a weight of over 80% in the latter.

Figure 1

The Canadian MCI evaluated at a relative weight of 3

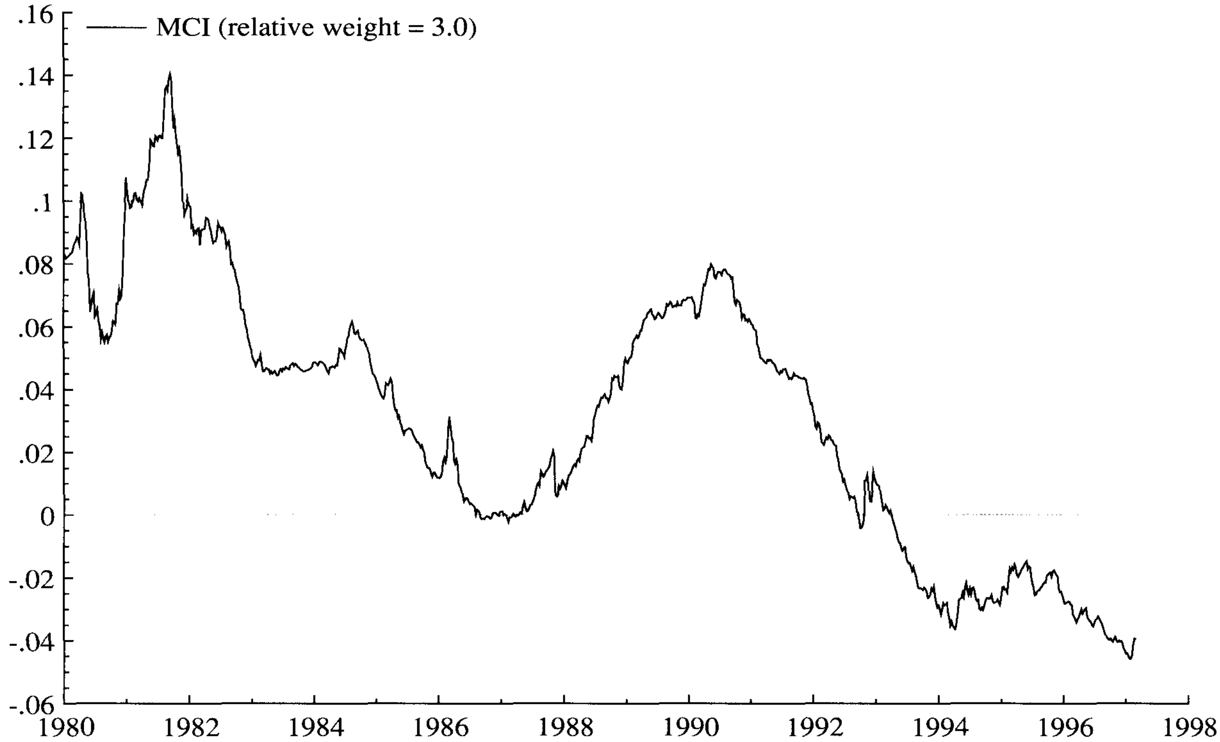
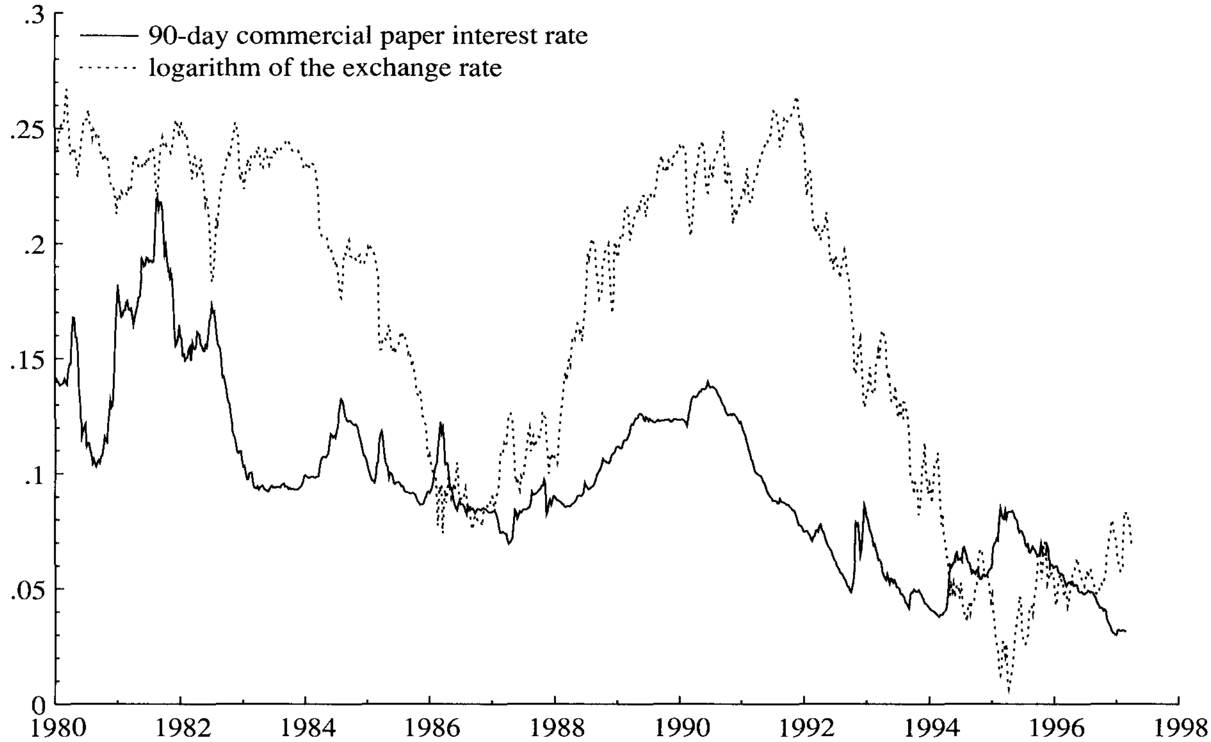


Figure 2

The components of the Canadian MCI: the 90-day commercial paper interest rate and the logarithm of the exchange rate



fall in the index is interpreted as a loosening of monetary conditions. As a policy indicator, the MCI aims to keep track of both interest rate and exchange rate movements and their effects on aggregate demand. From 1990 through 1993, the MCI fell steadily, signaling a general loosening of monetary conditions. In 1994 and early 1995, conditions tightened. Thereafter, the index resumed falling.

In Figure 1 and in all other figures of MCIs herein, each MCI is scaled such that its weights sum to unity, i.e., $\theta_R + \theta_e = 1$. A plotted MCI is thus always in units equivalent to the interest rate, measured as a fraction, thereby permitting easy interpretation of and comparison across different MCIs. For instance, the decline in the Canadian MCI from 1990 to 1994 is interpreted as the equivalent of a 12 percentage point (1,200 basis point) decline in the interest rate. Roughly half of this decline is due to the 20% depreciation of the Canadian dollar over that period, leading to Figure 2.

Figure 2 shows the two components of the MCI: the nominal 90-day commercial paper rate and the logarithm of the nominal G-10 trade-weighted Canadian dollar. During 1990 and 1991, the Canadian dollar remained relatively constant while the interest rate declined, with the latter variable being primarily responsible for the fall in the MCI. In 1992 and 1993, both variables moved downward, with both contributing to the MCI's continued fall. From 1994 onward, the two variables have moved in opposite directions, offsetting each other's movements to some extent.

2.2 The Bank of Canada's MCI as an operational target

The Bank of Canada has used its MCI as an operational target for several years. This subsection describes how the Bank has done so, focusing on the role played by the Bank's econometric model.

The Bank of Canada calculates a desired or target path for the MCI from interest rate and exchange rate forecasts from the Bank's Quarterly Projection Model (or QPM). The QPM includes equations for output growth (similar to (2)), inflation, and the exchange rate. In the model, interest rates and exchange rates influence output, which in turn influences inflation through a Phillips curve relationship. The exchange rate is determined through an uncovered interest rate parity condition with a risk premium. Additionally, the model incorporates a monetary response function, which is designed to bring inflation back to the midpoint of the Bank's inflation target range within a specified time, and subject to smoothness constraints on the path of the interest rate. Currently, the Bank has an inflation target range of 1 to 3% per annum at 6 to 8 quarters out. From the model, the Bank derives a solution for the future paths of the interest rate and exchange rate, consistent with the inflation target. The desired path for the MCI is then calculated from those paths on the interest rate and exchange rate.[3]

If, in the short term – from week to week – the actual MCI rises above (or falls below) its target path, this is interpreted as a tightening (or loosening) of monetary conditions relative to those anticipated and desired, and the Bank considers responding. In effect, the MCI is a convenient short-hand calculation for how to adjust interest rates if the exchange rate moves sometime between adjacent formal (quarterly) forecast rounds with the QPM. Operationally, at weekly and mid-quarter meetings, the MCI serves as a starting point in policy discussions, in which the Bank looks at developments that have occurred since the beginning of the quarter in deciding whether to adjust policy. The Bank then may also make adjustments to the desired path of the MCI.

Gordon Thiessen, Governor of the Bank of Canada, summarizes the role of the MCI at the Bank, as follows:

> ... we [at the Bank of Canada] aim at a path for monetary conditions that would bring about a path for aggregate demand and prices consistent with the control of inflation.
>
> Thiessen (1995, p. 54)

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[3] In practice, the Bank controls the overnight interest rate, which is closely linked to the 90-day commercial paper interest rate. For the most part, the discussion below ignores the distinction between the Bank's actual policy instrument and what constitutes a very short-term operational target.

Charles Freedman, Deputy Governor of the Bank of Canada, provides additional details:

> In the last few years, the Bank of Canada has used the concept of monetary conditions (the combination of the movement of interest rates and the exchange rate) as the operational target of policy, in much the same way as short-term interest rates were used in the past.
>
> ... The objective of monetary policy over the next three years or so is to maintain the rate of inflation within a band of 1 to 3 per cent. The quarterly Bank of Canada staff projection takes into account such factors as the movements in foreign variables and domestic exogenous variables as well as the momentum of the economy, and sets out a path for monetary conditions that will result in the rate of inflation six to eight quarters ahead being within the Bank's target band. ... One can think of this path [of the MCI] as the desired or target path for monetary conditions.
>
> Freedman (1995, pp. 53, 54, 56)

Three qualifications should be noted. First, while the QPM is the foundation for generating the forecasts, additional analyses of the domestic and foreign economies also play a role, with iterations between sectoral specialists and the QPM resulting in judgmentally adjusted forecasts. Second, the forecasts are conditional, both on the Bank's views of future domestic monetary and fiscal policy and on its views of future foreign economic outcomes. Third, there are operational considerations, as described in the next subsection.

2.3 Operational considerations

Implementing an MCI as a target involves practical, operational considerations, which are reflected by the Bank of Canada's experience. These considerations include both the timing of policy adjustments and the role of additional information in the policy process. Timing, or "tactical" considerations, has sometimes made an MCI a difficult operational target to achieve.

The Bank has a desired path for the MCI. If the actual MCI is "off course", then the Bank tries to move it back on track as quickly as is tactically possible. "Tactically" is the operational word here, in that the Bank sometimes has allowed actual and desired MCIs to differ for considerable periods – of a quarter or more. The Bank has explained such episodes by arguing, for example, that observed exchange rates were out of line relative to fundamentals, as the Bank believed happened with transitory reactions to the Quebec problem. In such situations, the calculated index may not accurately reflect intended or actual monetary conditions. Freedman (1995) provides a lucid account of this problem:

> ... Suppose an easing of monetary conditions was appropriate, but there was a great deal of uncertainty and nervousness in the exchange market. ... In such circumstances, the Bank would delay any decision to ease monetary conditions because of the risk that an action to reduce the overnight rate could result in significant weakness in the exchange market and lead to the buildup of extrapolative expectations in that market, followed, as we have so often seen in Canada in recent years, by an increase in interest rates in the money market and the bond market. In effect, an attempt to ease monetary conditions could, via the interaction of developments in the exchange market and domestic financial markets, result in an outcome where monetary conditions ended up tighter and not easier. Thus, the tactical aspect involves choosing the timing of changes to avoid undesired market-driven outcomes. (p. 58)

Timing is clearly an issue. Furthermore, market conditions and the market's responses to Bank actions may simply prevent the Bank from achieving its target, at least in the short or medium term.

The MCI is central to the Bank's decision process, in which the use of the MCI is viewed as interest-rate targeting, adjusted for exchange-rate effects on aggregate demand in a small open economy. That said, inputs additional to the MCI do influence the Bank's policy decisions, as Freedman (1995) indicates:

> ... While the [model-based path for the MCI] recommended by the staff is a crucial input into the views of senior management on the desired path for monetary conditions, senior management may also incorporate into its thinking the possible effects of a broad range of outcomes with respect to the movements of exogenous variables or the momentum of the Canadian economy. Indeed, the staff prepares alternative "risk scenarios" that incorporate some of these factors. Management may also decide in which direction to take or avoid risks (e.g., that it is appropriate to be especially vigilant about a resurgence of inflation). If, following this type of analysis, there is a divergence between actual and desired monetary conditions, the Bank will look for the right time to make adjustments. Among the factors that enter into the timing decision are market uncertainty and market nervousness. (p. 59)

Thus, even for a stable developed economy like Canada, achieving targeted levels of the MCI has sometimes proven infeasible because of tactical difficulties. For countries with much more volatile economies and larger speculative swings in the exchange rate, tactical considerations are even more likely to make an MCI operationally infeasible.

2.4 General usage of MCIs

While much discussion in the literature focuses on the Bank of Canada's use of its Monetary Conditions Index, MCIs have widespread use among other institutions and for other countries. The central banks of New Zealand, Norway, and Sweden each have published an MCI and (to varying degrees) use it in conducting monetary policy. The Reserve Bank of New Zealand (starting in late 1996) uses an MCI as an operational target in much the way that the Bank of Canada does; see the Reserve Bank of New Zealand (1996). The central banks of Norway and Sweden use MCIs in a more limited fashion – as indicators of monetary conditions when formulating their monetary policies; see Norges Bank (1995) and Hansson and Lindberg (1994). In a recent paper, Dornbusch, Favero, and Giavazzi (1998) construct an MCI for the European Central Bank over a region spanning most of the European Monetary Union (EMU). The IMF and the OECD also use MCIs in evaluating monetary policies across countries; and businesses such as Deutsche Bank, Goldman Sachs, JP Morgan, and Merrill Lynch calculate MCIs to evaluate different countries' monetary conditions.

Table 1 compiles alternative relative weights for MCIs across selected countries, as published or made available by the institutions and authors just mentioned. This table is indicative of the range of countries and sources, rather than being exhaustive. While MCIs are about *monetary* conditions, institutions other than the central bank of a given country may well calculate an MCI for that country, even if that central bank does not publish or use an MCI in policy. For many countries, several estimates of the relative weights are available, and the estimates vary considerably. In light of the range in available weights, Section 3 considers inter alia the empirical consequences of using different weights.

The range of estimated weights in part reflects the use of different models and different sample periods. However, a given range of estimated weights across a set of models and sample periods has no implications for the confidence intervals of any model's estimated relative weight, not even for those of a correctly specified model's estimated relative weight. A consensus in estimated weights across models would reflect just that – a consensus – and nothing more. For instance, Freedman (1994, pp. 469-70) reports similar estimates of relative weights across a range of Canadian models. That consensus implies nothing about confidence intervals for those estimated weights. Such a consensus could easily arise if the different models of a given economy used more or less the same data: specifically, the different models' estimated relative weights are unlikely to represent *independent* random draws on some unknown relative weight. Section 4 thus examines the uncertainty of the estimated weights and the empirical consequences that such uncertainty has for using an MCI as an indicator or target.

Table 1

Selected alternative relative weights for MCIs

Country	Source							
	Central banks	**IMF**	**OECD**	**Deutsche Bank**	**Goldman Sachs**	**JP Morgan**	**Merrill Lynch**	**Dornbusch et al.**
Australia			2.3			4.3	4	
Austria				3.3				
Belgium						0.4		
Canada	2, 3	4, 3	2.3		4.3	2.7	3	
Denmark						1.9		
EMU								2.17
Finland						2.5		
France		3	4	3.4	2.1	3.5		2.10
Germany		2.5, 4	4	2.6	4.2	2.3	4	1.39
Italy		3	4	6.6	6	4.1		2.89
Japan		10	4		8.8	7.9	10	
Netherlands				3.7		0.8		
New Zealand	2							
Norway	3					1.4		
Spain			1.5	2.5		4.2		1.46
Sweden	3–4		1.5	0.5		2.1		8.13
Switzerland				6.4		1.7		
United Kingdom		3	4	14.4	5	2.9	3	
United States		10	9		39	10.1	10	

Notes: Weights are those on interest rates relative to those on exchange rates.

Sources: Bank of Canada (1995, p. 14), Reserve Bank of New Zealand (1996, pp. 22–3), Norges Bank (1995), Hansson and Lindberg (1994, p. 16), International Monetary Fund (1996a, p. 16; 1996b, p. 19; 1997, p. 24), OECD (1996, p. 31), Gräf and Schonebeck (1996), Davies and Simpson (1996), Suttle (1996), Merrill Lynch (1997), and Dornbusch, Favero, and Giavazzi (1998, Table 5.6), with additional information on specific MCI weights for Deutsche Bank, Goldman Sachs, and JP Morgan from personal communications with Theodor Schonebeck, John Simpson, and Carl Strong.

3. Two facets of MCI design

This section analyzes two facets in the design of an MCI: the choice of weights and variables (Section 3.1), and the assumptions of the underlying empirical model (Section 3.2), with the latter leading directly into Section 4 on coefficient uncertainty. Empirically, MCIs appear very sensitive to even minor changes in weights, variables, and assumptions.

3.1 Choice of weights and variables

The choice of weights and variables in an MCI is central to constructing the index itself, and MCIs can be empirically sensitive to that choice. For example, Figure 3 plots three alternative MCIs for Canada: one using the Bank of Canada's relative weight of 3, and the other two using the smallest and largest Canadian weights appearing in Table 1 (weights of 2 and 4.3 respectively). In 1986 and 1987, the MCI is nearly invariant to the relative weight because the exchange rate was virtually constant; see Figure 2. From 1988 through 1994, weights matter considerably because the Canadian dollar appreciated and then depreciated. In late 1996, some versions of the index actually move in different directions. To focus on this phenomenon, Figure 4 plots the MCIs from the

Figure 3

The Canadian MCI evaluated at relative weights of 3, 2 and 4.3

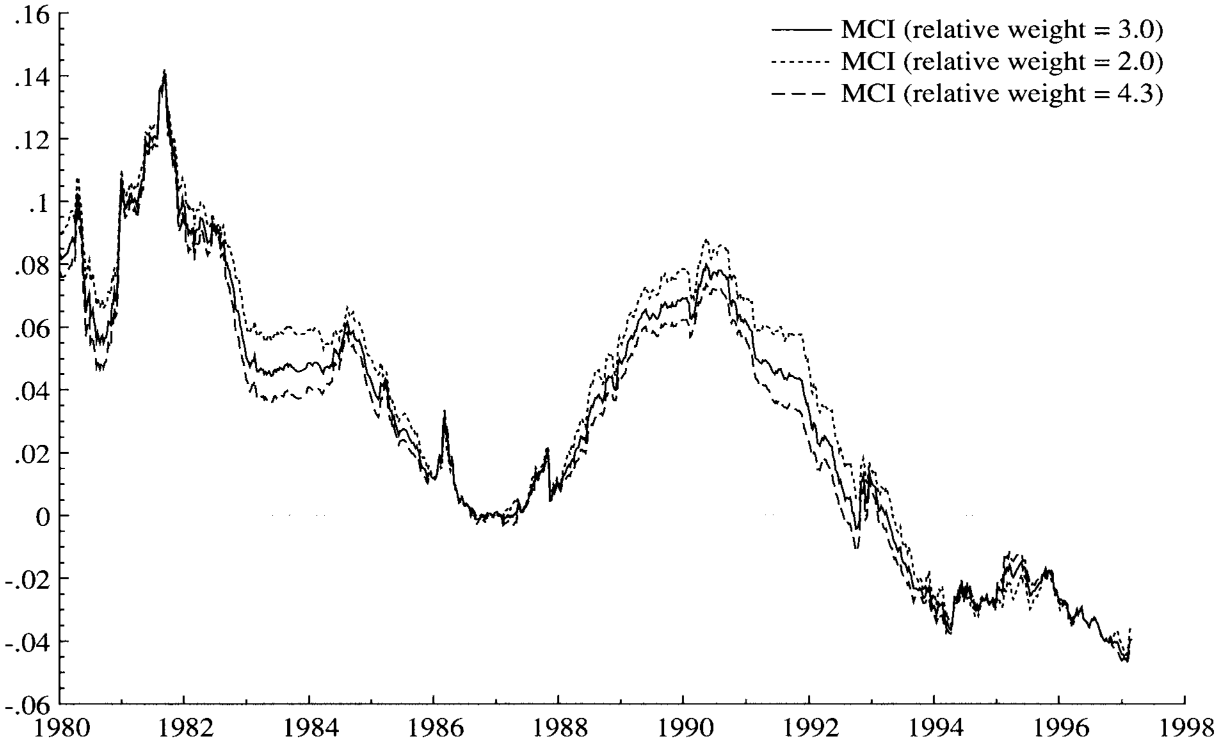
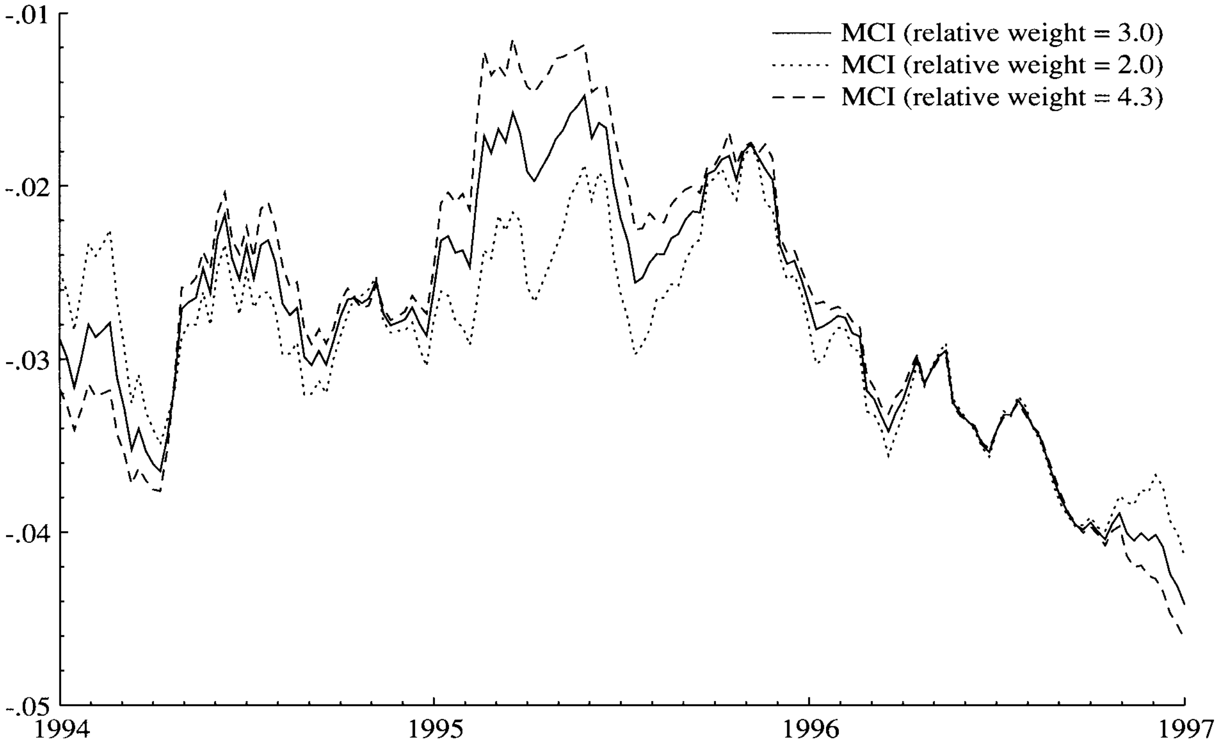


Figure 4

The Canadian MCI evaluated at relative weights of 3, 2 and 4.3 over a recent subsample



beginning of 1994 onwards. The MCI with the smallest relative weight indicates a moderate tightening of monetary policy in the autumn of 1996, whereas the indexes with larger weights on the interest rate show a moderate or considerable loosening. Similar episodes occur in 1983–84, while noting that the extreme scale of Figure 3 necessary for capturing 17 years of data does dwarf the discrepancies in the various MCIs' behavior. In general, the weights can and do affect the magnitude and the sign of changes in the index. Notably, the Bank of Canada initially used weights of 2:1 and then switched to 3:1, with a substantial effect on the measured MCI.

The selection of variables in the MCI is an open issue as well. MCIs in this paper's figures are calculated from only a single interest rate and a single exchange rate. Many possible interest rates and exchange rates are available, and using different variables in the MCI can induce differences in movement similar to those encountered in the choice of weights.

For example, the Bank of Canada currently uses a nominal G-10 bilateral trade-weighted exchange rate. Such an exchange rate is appealing in trade equations, and hence in an aggregate output equation such as (2). However, for short-term to medium-term monetary policy, international exchange markets may be more speculative and financial in nature, in which case bilateral *trade* weights are less germane. A conflict may well exist between the data appropriate for the underlying econometric model and those appropriate for policy analysis.

A range of alternative interest rates also exists. Numerous short-term rates are available, and MCIs need not be restricted to short-term rates alone. Deutsche Bank and Goldman Sachs in particular use weighted averages of long-term and short-term interest rates in calculating their MCIs. Furthermore, an MCI may be calculated from real (rather than nominal) variables, with the measurement of expected inflation implying yet an additional choice in variables. Without substantially better information on the choice of weights and variables, currently calculated MCIs may well be misleading about underlying changes in monetary stance, both in magnitude and in sign.

The choice of variables also can be viewed as an issue in aggregation, with four forms of aggregation occurring. First, an MCI includes only exchange rates and interest rates, to the exclusion of other potential variables; see Section 3.2 below. That exclusion constitutes aggregation, with weights of zero on those other variables. Second, bilateral exchange rates are aggregated into a single exchange rate index. Third, available interest rates are aggregated, often into a single interest rate. Fourth, combining a given exchange rate index and a given interest rate into an MCI constitutes aggregation. All four senses of aggregation involve losses of information, and use of an MCI implicitly assumes that the information lost is not important for policy. Specifically, because many combinations of an interest rate and an exchange rate give rise to the same value of the index, the lost information is important if the mix of the two variables is of concern.

In practice, the Bank of Canada does look at the two variables separately, especially when considering how to alter monetary conditions; see Freedman (1995, p. 58), as quoted above. For example, rapid depreciation of the exchange rate can translate into risk premia across the yield spectrum, in which case the exchange rate may be weaker in conjunction with higher interest rates, but overall monetary conditions can be unaffected. As Figures 1-4 show, episodes exist (such as early 1994) when the interest rate increased but the exchange rate weakened even more, with the MCI moving in the *opposite* direction from the interest rate.

3.2 Assumptions of the underlying empirical model

The use and interpretation of an MCI rest upon the assumptions of the underlying model. Several issues arise for that model, including dynamics, data nonstationarity and differencing, exogeneity and feedback, parameter constancy, the choice of model variables, and the uncertainty arising from estimating the model. This subsection summarizes these six issues, relating them to the corresponding model assumptions. These assumptions are often testable and, if violated, directly affect the economic interpretation of the MCI. Eika, Ericsson, and Nymoen (1996a) present a more detailed analytical assessment, and their empirical evaluation confirms such difficulties in models for the Canadian, Swedish, and Norwegian MCIs. Gerlach and Smets (1996) and Alexander (1997) discuss possible economic theoretical underpinnings of an MCI and the associated, rather stringent assumptions required for an MCI to be an optimal policy target.

First, the relationships between the policy instruments, the exchange rate, the short-term interest rate, output, and inflation generally are *dynamic*, implying different short-run, medium-run, and long-run multipliers. Thus, the policy horizon may affect the relative weight. If policy is concerned with several horizons, the weight for a single horizon may not be adequate.

Second, the temporal properties of the data themselves bear on the construction of an MCI. In particular, *nonstationarity* of the data (e.g., as in a series with drift) may affect the distribution of the error terms in the associated model and thereby affect statistical inference. Nonstationary data also may be cointegrated. If so, the relevant equations should include levels of the series, and calculations of multipliers should account for those levels. By contrast, output equations for calculating MCI weights are typically estimated with differenced or detrended data, with no testing for cointegration. Furthermore, the MCI itself is calculated on the levels of the data. Adjustment of the MCI relative to a base period simply subtracts a constant from an unbased MCI and does not constitute working with differenced data. The mixed use of differences and levels affects the interpretation of the weights: short run for differences, contrasting with long run for levels.

Third, the postulated *exogeneity* of the policy instruments and other variables is potentially misleading. In the MCI itself, the weights are interpreted as elasticities of aggregate demand with respect to the interest rate and the exchange rate. This interpretation assumes no feedback from aggregate demand or inflation onto exchange rates and interest rates over the relevant policy horizon. Such feedback may occur under any policy regime and seems likely to occur under inflation targeting by design. With feedback, the estimated weights need not reflect the total effects of the exchange rate and interest rate on aggregate demand. As an alternative, the feedback could be estimated and subsequently incorporated into the elasticities from which an MCI is derived.

Fourth, *parameter constancy* is critical to the interpretation of an MCI, and it turns on all three of the aforementioned issues. Statistically nonconstant weights may arise empirically from misspecified dynamics, improper treatment of nonstationarity, or incorrect exogeneity assumptions. Because the MCI is designed for policy, it is important to establish the invariance of the weights to changes in policy, yet this conjectured invariance generally has not been investigated empirically. With nonconstant parameters, estimation over different sample periods would result in different estimates of the weights, and so different choices of weights. Eika, Ericsson, and Nymoen (1996a, Section IV) illustrate how that choice of weights can affect policy inferences with an MCI.

Fifth, the *choice of model variables* determines the variables omitted from the model. Significant omitted variables in the model's relationships may affect dynamics, cointegration, exogeneity, and parameter constancy in the model.

More generally, the use and interpretation of an MCI in policy assumes the existence of direct and unequivocal relationships between the variables involved. Possible additional influences in those relationships can confound the strict interpretation of an MCI as an index of monetary conditions.

One such relationship is that between the actual policy instrument (such as the central bank's overnight interest rate) and the exchange rate and short-term interest rate. If variables other than the policy instrument play an important role in determining the exchange rate and interest rate, neglect of those other variables has substantive implications for policy with an MCI. For example, changes in world oil and commodity prices may alter a country's terms of trade, thereby affecting the exchange rate. The MCI would then change, even if monetary stance remained unchanged. Likewise, changes in world interest rates and inflation rates and changes in domestic asset portfolio preferences may alter the domestic short-term interest rate, and so the MCI. The variables from which the MCI is constructed may reflect phenomena other than just direct monetary policy, so movements in the MCI are not tied unequivocally to changes in monetary stance. Conversely, by following or targeting an MCI, a central bank could be misled into adopting an overly tight or loose monetary policy, simply because some external shock affected the exchange rate or the domestic short-term interest rate.

An additional relationship is the one between exchange rates, interest rates, and output growth, which is the basis for calculating the relative weight in the MCI. Exchange rates have other effects on the economy, such as influencing domestic prices directly; cf. Froot and Rogoff (1995), Juselius (1992), and de Brouwer and Ericsson (1998) for examples of theoretical and empirical research supporting such a channel. The interest rate also may have other channels to inflation. Specifically, interest rates may affect mortgage payments and hence inflation through the calculated cost of housing. Neglect of these transmission mechanisms is likely to result in the MCI being a misleading index of monetary conditions per se, particularly if the MCI is being used by the central bank in targeting inflation. An MCI focuses on only one of many potential channels and on only two of many potential variables in the monetary transmission mechanism.

Sixth, the relative weight in an MCI is based on an estimated empirical model, and so is subject to coefficient *uncertainty* from that estimation. Thus, the estimated weight may be numerically nonconstant, even if it is statistically constant. Numerically nonconstant weights may arise from the lack of information content in the data, leading to large standard errors. Section 3.1 above shows that the calculation of an MCI can be sensitive to the choice of weights. Uncertainty from estimation has not been previously examined for MCIs, so the next section (Section 4) turns to quantifying that uncertainty and assessing its consequences for using an MCI in practice.

4. The uncertainty of MCI weights and some consequences

This section assesses the statistical uncertainty from estimating the MCI weights in models for Canada, New Zealand, Norway, Sweden, and the United States. This section then summarizes the policy consequences of uncertainty in an MCI's relative weight and considers some alternatives. See Ericsson, Jansen, Kerbeshian, and Nymoen (1997) for details on the statistical framework employed.

To assess the uncertainty of an MCI weight, an equation is estimated in the form:

$$\Delta y_t = a\Delta RR_t + b\Delta q_t + other \;\; variables + error . \tag{3}$$

The relative MCI weight μ is a/b, and its estimated value $\hat{\mu}$ is $\hat{a}/\hat{b}$, where a circumflex denotes estimation. Confidence intervals for the estimated MCI weight can be constructed from a Wald statistic, a likelihood ratio statistic, or a Fieller statistic inter alia; see Wald (1943), Silvey (1975, pp. 115-8), Fieller (1940, 1954), and Kendall and Stuart (1973, pp. 130-2). As Gregory and Veall (1985) discuss in general and Ericsson, Jansen, Kerbeshian, and Nymoen (1997) discuss for the MCI relative weight in particular, the likelihood ratio approach has distinct advantages over the other two approaches, so it is used below.[4] Similar issues arise in calculating the estimated uncertainty of NAIRUs; see Staiger, Stock, and Watson (1997).

Table 2 lists estimated MCI relative weights and their confidence intervals from models for Canada, New Zealand, Norway, Sweden, and the United States. Confidence intervals are calculated for 95%, 90%, and 67.5% (i.e., "±1 standard error") levels. For each country, the interest rate is measured as a fraction and the exchange rate is in logarithms. The models are taken from Duguay (1994, p. 50, Table 1, column 7), Dennis (1997, p. 14, Table 2, Equation A), Jore (1994, Equation 2), Hansson (1993), and Ericsson, Jansen, Kerbeshian, and Nymoen (1997) respectively. See also Eika, Ericsson, and Nymoen (1996a, 1996b) and Hansson and Lindberg (1994) for additional analysis of the Norwegian and Swedish models. Following the various central banks' practices, the Canadian MCI is nominal, whereas those for New Zealand, Norway, and Sweden are real. The Federal Reserve Board does not publish an MCI, so the choice of a nominal MCI versus a real MCI is open for

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4 That said, the confidence intervals for the estimated MCI weights in Dornbusch, Favero, and Giavazzi (1998, Table 5.6) are calculated using the Wald statistic.

the United States. Below, the nominal MCI is used for the United States. Ericsson, Jansen, Kerbeshian, and Nymoen (1997) provide additional details on all five models.

Table 2

MCI relative weights and their estimated confidence intervals

Calculation	**Country**				
	Canada	**New Zealand**	**Norway**	**Sweden**	**United States**
MCI relative weight					
Published	3	2	3	3–4	–
Estimated	3.56	1.75	2.15	2.02	–3.69
Confidence interval					
95% level	[0.74, ∞]	[0.30, 7.31]	[0.00, ∞]	[1.06, 2.96]	[–∞, ∞]
90% level	[1.06, ∞]	[0.52, 5.05]	[0.36, 26.6]	[1.27, 2.76]	[–∞, ∞]
67.5% level	[1.80, 9.60]	[0.97, 3.04]	[1.00, 4.98]	[1.61, 2.43]	[–8.45, 1.84]

Notes: The published MCI relative weights are those used by the corresponding central banks; see Table 1. The estimated MCI relative weights are calculated for a long-run horizon from the models reported in Ericsson, Jansen, Kerbeshian, and Nymoen (1997). The estimated confidence intervals are constructed from likelihood ratio statistics for those models at the reported significance levels.

For Canada, the estimated relative MCI weight is 3.56, somewhat larger numerically than the estimate of 2.67 from Duguay's equation, but well within the range of estimates typically obtained for Canada (see Table 1). The 95% confidence interval is enormous: [0.74, ∞]. It includes equal weights on the interest rate and exchange rate ($\mu = 1$) as well as a zero weight on the exchange rate ($\mu = \infty$). Even the 67.5% confidence interval, equivalent to a plus-or-minus one standard error band for single coefficient estimates, is large: [1.80, 9.60]. This high degree of uncertainty is unsurprising, given the marginally significant coefficient on the exchange rate in (2).

For New Zealand, the estimated relative MCI weight is 1.75, and the 95% confidence interval is [0.30, 7.31]. While the confidence intervals for New Zealand are not as large as those for Canada, the presence of small relative weights in the confidence interval can have a marked effect on the calculated MCIs.

For Norway, the estimated relative MCI weight is 2.15, and the 95% confidence interval is [0.00, ∞], even larger than those calculated for Canada and New Zealand. All non-negative weights fall within the 95% confidence interval, and completely different accounts of monetary conditions are feasible with different empirically acceptable estimates of the MCI relative weight.

For Sweden, the estimated relative MCI weight is 2.02, and the 95% confidence interval is quite small: [1.06, 2.96]. However, the calculated confidence intervals are probably unreliable, given that the over-identifying restrictions in the Swedish model are rejected against the corresponding unrestricted reduced form; see Eika, Ericsson, and Nymoen (1996a, Appendix). Even if a 95% confidence interval of [1.06, 2.96] is assumed, MCIs still can differ by a few hundred basis points, depending upon which value of the relative weight is chosen from that interval.

Although the Federal Reserve Board does not publish an MCI for the United States, other institutions do, as Table 1 shows. To calculate the uncertainty of an estimated MCI weight for the United States, we estimate a model for the growth rate of real US GDP that is similar in form to the Canadian and Norwegian models. The real interest rate has a negative coefficient whereas the real exchange rate has a positive coefficient, so the estimated relative MCI weight is negative: –3.69, numerically. A negative coefficient such as this is difficult to interpret economically. However, the 95% confidence interval is the entire real line, [–∞, ∞]. Even the 67.5% confidence interval is large: [–8.45, 1.84].

For the models examined, the large estimation uncertainty associated with the MCI weights renders calculated MCIs uninformative for policy. The model for the Swedish MCI is itself misspecified, so it is difficult to interpret MCIs based on that model. For all five countries, the estimation uncertainty often implies discrepancies in the calculated MCI of 100 basis points or more for statistically acceptable choices of the relative weight. Such discrepancies occur even at one quarter ahead, which is a short horizon for policy based on an MCI. Furthermore, the choice of weight often affects the MCI's direction of movement, and not just the magnitude of its movement. That feature is particularly problematic, in so far as an MCI is interpreted as an *indicator* of monetary stance. See Ericsson, Jansen, Kerbeshian, and Nymoen (1997) for further graphical evidence of the marked numerical consequences on the MCIs from estimation uncertainty.

These results on estimation uncertainty could not have been known a priori: the confidence intervals could have been small, but they were not. Confidence intervals could still be small for relative MCI weights derived from *other* models of these countries' economies, or from models of other countries' economies. However, for the models studied, which were developed at the countries' respective central banks, these confidence intervals are large empirically.

While MCIs as such appear impractical for use in policy, their motivation is sensible. In a small open economy, foreign economic activity is likely to affect the domestic economy through the exchange rate; and empirical models are a potentially sensible way of capturing the exchange rate's effects.

That said, tools other than MCIs are available for policy input in this context. In the conduct of monetary policy, central banks historically have considered a wide range of economic variables, including but not limited to interest rates and exchange rates. Central banks have changed their emphasis across those variables over time, for instance, in light of financial innovation. Instead of summarizing model-based calculations in a single index such as an MCI, those calculations may be presented directly, as time-dependent effects across a variety of economic aggregates. Such model-based calculations also may then be part of the economic information feeding into the policy process itself.

Policy-oriented examples of model-based calculations for Canada, Norway, and the United States appear in Poloz, Rose, and Tetlow (1994), Norges Bank (1996), and Mauskopf (1990). Good graphical and tabular techniques can ease the burden in communicating inherently multivariate results; see Tufte (1983, 1990, 1997). Furthermore, better design of empirical models for policy appears possible, using econometric tools and corresponding software developed over the last decade or so. Spanos (1986), Banerjee, Dolado, Galbraith, and Hendry (1993), and Hendry (1995) describe some of those tools; Doornik and Hendry (1996) exemplifies the software available; and the papers in Ericsson and Irons (1994) inter alia show how such tools and software can aid empirical modeling.

Conclusions

An MCI is an appealing operational target for monetary policy – it broadens an interest-rate target to include effects of the exchange rate on an open economy. In doing so, an MCI also incorporates model-based estimates of the effects of monetary policy on the economy. Notwithstanding the intuitive attraction of a Monetary Conditions Index, substantive limitations in the index's use arise from tactical difficulties, the choice of weights and variables, the underlying model's assumptions, and the associated uncertainty of the estimated relative weight. The latter three issues pertain to summary indicators and model-based calculations generally, but they appear particularly important empirically for MCIs. As a policy target and as an indicator of monetary conditions, an MCI focuses on only two of many potential variables in the monetary transmission mechanism. While the Bank of Canada and the Reserve Bank of New Zealand currently use MCIs as operational targets, they are well aware of the shortcomings involved. This paper has reviewed and interpreted the use of an MCI, focusing on the implications of estimation uncertainty for the practical implementation of MCIs in monetary policy.

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Comments on "Understanding a monetary conditions index" by Neil R. Ericsson, Eilev S. Jansen, Neva A. Kerbeshian and Ragnar Nymoen

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by Dinah Maclean

This paper provides a thorough analysis of problems associated with MCIs, in particular the confidence bands around the MCI weights. As stressed by the authors, it is very important to be aware of the uncertainty surrounding these weights, and to understand the sensitivity of results to different values. Nevertheless, I disagree with the paper's conclusion that calculated MCIs are impractical for use in policy. Quite the reverse: used correctly, the MCI is a very useful tool between projections, which serves as a reminder that both interest rate and exchange rate movements have important affects on monetary conditions.

The words "used correctly" are very important. Many of the criticisms of the MCI, including many of those in this paper, are not appropriate given the way it is used within the Bank of Canada and elsewhere. Although the paper does have a good description of the manner in which it is used, given the importance of this issue it is worth highlighting some key features.

Four times a year the staff undertake a full projection. Ericsson et al. state that, instead of summarizing model-based calculations in a single index such as the MCI, these calculations should be presented directly so that they are part of the economic information feeding into the policy process. But this is already the case. At each projection, the staff try to get as complete an accounting of all the shocks and factors affecting the economy as possible. The judgement of those people who monitor the day-to-day data releases and provide short-term outlooks is combined with the output of a general equilibrium model which provides projections of a wide range of variables out into the medium term.

The presentation of the projection to senior management includes a full discussion of the main factors underlying the outlook, provision of alternative scenarios which consider how things would differ given slightly different assumptions, and of course a full range of tables including, but not limited to, the MCI.

As the authors recommend, therefore, model-based calculations are already a key part of the policy process. It is not practical, however, to run a projection every week, or even every month. In between times, the staff need other means to judge the implications of changing conditions.

One of the tools used between projections is the MCI. The MCI provides a benchmark for monetary conditions, consistent with the starting point assumptions in the projection, and consistent with the achievement of the inflation target. As new information becomes available, the MCI provides a guideline for comparison. It is not, however, taken as a literal track which must be adhered to. It is also not the only thing which is considered.

Ericsson et al. rightly state that it is important to assess the source of changes in interest and exchange rate movements before deciding how to respond. This is carried out on an on-going basis by the staff. If, for example, movement in the exchange rate is linked to a change in commodity prices, clearly the desired monetary conditions will alter. If the shock is large enough, the staff may update the last projection, including the new shock. If, on the other hand, exchange movements seem to be driven by changes in confidence or portfolio adjustment, as they frequently are for Canada, the desired monetary conditions will not alter. Ignoring the impact of, for example, an appreciation of the exchange rate caused by such factors, and the resulting tightening of monetary conditions, would lead to the risk of a significant policy error.

As the authors point out, tactical considerations may mean that it is hard or undesirable to reach the desired MCI at a particular point in time. Nevertheless, a similar argument may be made about short-term interest rates, or any other short-term operational target. Moreover, the fact that it may not always be possible to reach it, does not mean one should not have a desired path in mind.

The authors criticise the choice of variables within the MCI, particularly the fact that it is limited to two. But it is hardly an arbitrary selection. In an open economy, with flexible exchange rates, monetary policy works through both interest rates and the exchange rate. Certainly, however, there are many different measures of the interest rate and the exchange rate that could be used.

That there is great uncertainty about the weights is clearly true. In fact, it seems unlikely that the weights will be stable over time, or over a cycle. With the increase in Canada's propensity to import over the last decade, particularly in interest rate sensitive goods, for example, it is quite likely that exchange rate movements have relatively more impact than previously. It is not surprising, therefore, if reduced form estimates are highly sensitive to the sample period chosen.

For this very reason, the weights for the MCI used by the Bank of Canada were chosen by looking at a wide range of information sources including reduced form equations and estimated VARs. In general, the range of estimates suggested relative weights between 2:1 and 4:1. The choice of 3:1 was, therefore, one where the risks were seen as relatively balanced.

In addition, the weights are chosen as being relevant to a very specific time horizon: they are interpreted as being the relative movements in the exchange rate and interest rates which will leave output unchanged two years ahead. In other words, the current weights suggest that if the exchange rate depreciates by 1 percentage point, an increase in short-term interest rates one-third of the size will offset the impact on output, eight quarters ahead.

Given the way in which the weights were chosen, therefore, it is unfair to judge them purely by the standard of the confidence bounds around a reduced form estimate. Moreover, while there is uncertainty over the coefficient values, the coefficient on the exchange rate for Canada is clearly not zero. Thus ignoring exchange rate movements and considering only the interest rate would not be appropriate.

In conclusion, while the paper explores what are clearly important issues, many of the criticisms are only relevant if the MCI is used as a rigorous rule, which is not the case. The MCI is used as a benchmark, consistent with the view of the world when the projection was undertaken. It provides a link between projections, but does not replace them, or narrow their focus. It is a tool to help in the analysis of shocks and their implications, not a substitute for such analysis. Above all, the MCI serves as a reminder that both interest rates and exchange rates are important to monetary conditions.

The credit channel in the transmission of monetary policy: the case of Spain

Ignacio Hernando*

Introduction

The existence of a credit channel in the process of transmission of monetary impulses is a recurrent topic in the literature on the effects of monetary policy. Nevertheless, macroeconomic models have frequently tended to ignore this channel of transmission and empirical tests of its existence have usually been unsatisfactory. Over the past decade, a large number of papers, mainly about the US economy, have helped to correct the traditional assumption that credit is relatively unimportant in the monetary transmission mechanism. They have also provided new empirical approaches to test the existence of this credit channel.

Conventional wisdom regarding the effects of monetary policy on the real economy has concentrated on the so-called money channel, using models based on two financial assets (money and bonds). Under this approach, in the event of an alteration in the supply of money in a given economy, equilibrium will be restored by changes in the interest rate for bonds that will ultimately have an effect on real variables. This approach implicitly assumes that there is only one alternative asset to money (or that all alternative assets are perfect substitutes) and may therefore reflect a partial view of the monetary transmission mechanism. Such a partial approach could lead to erroneous conclusions regarding the degree of effectiveness of monetary policy or the usefulness of the different variables as intermediate targets.

An exhaustive study of how monetary impulses are transmitted requires consideration of the credit channel. The literature uses this term to refer to two different, albeit related, transmission processes. First, in accordance with the bank lending channel (or credit channel in the strict sense) – represented in models based on three assets: money, bonds and bank loans (see, for instance, Bernanke and Blinder (1988)) – the effects of monetary policy are not just via its impact on interest rates for open market transactions but also the result of its independent impact on the supply of bank loans. Secondly, the balance-sheet channel (or credit channel in the broad sense) refers to the additional effects of monetary policy on the final variables through variations in the net financial income received by agents and on their net wealth. Based on more extensive studies by Bernanke (1993), Kashyap and Stein (1994) and Gertler and Gilchrist (1993), this paper describes the theoretical grounds supporting the existence of the credit channel, distinguishing between the two senses of the term.

Consideration of the credit channel contributes to a fuller understanding of the response of output and of demand components to monetary policy measures. Kashyap and Stein (1994), for instance, list a number of reasons why it is important to take the existence of a credit channel into account. One is that the credit channel supports the existence of a different impact of monetary policy depending on agents' degree of access to capital markets. For their part, van Ees et al. (1994) present a model – comprising two types of firm, local and international, with the former lacking access to

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public capital markets – where the credit channel may even play an effective role in a small, open economy with fixed exchange rates and high international mobility of capital.

The credit channel should not be envisaged as an isolated, independent path, but rather as just another component in the complex workings of the mechanism for monetary policy transmission. As Bernanke and Gertler (1995) point out, the credit channel should be considered a mechanism that reinforces the traditional money channel. How important it is depends on various factors conditioning the relationship between the monetary authority's interventions and economic agents' expenditure decisions. Two such factors that are worth noting are, on the one hand, the institutional framework governing the financial system and, on the other, economic agents' financial position. Regarding the former, the extent to which the banking system acquires liabilities that are not subject to reserve requirements and the role played by non-bank intermediaries are key determinants of the relative importance of the credit channel. Related to this point, there are also implications for the implementation of the single monetary policy in the European Union. Along this line, Dornbusch et al. (1997) suggest that monetary policy shocks could have different effects across member countries depending on their financial structures. Regarding the second above-mentioned factor, as Peñalosa (1996) analyses in some detail, economic agents' financial position has a decisive influence on the transmission of monetary impulses. In this sense, the financial health of economic agents, insofar as it determines their degree of access to credit and the terms under which it is granted, affects the relative importance of the credit channel.

The aim of this study is to provide evidence, in the Spanish case, regarding the existence of the credit channel in the transmission of monetary disturbances. Unlike most of the empirical approaches to this topic, which are based on analysis of the response of money, credit and output to monetary impulses, this paper follows the Kashyap et al. (1993) approach and examines the relative behaviour of bank loans and some alternative source of financing for economic agents. For instance, Table 1 illustrates the considerable changes observed in the financial structure of companies. Such variation could be explained, at least partly, by changes in the extent and terms of access to different sources of finance induced by shifts in monetary policy. It is, moreover, expected that the composition of the financing also has a bearing on the determination of the real variables. With these considerations in mind, the test presented here, which is based on Kashyap et al., takes as a reference point the composition of corporate liabilities. The basic idea is that, when the credit channel exists, monetary policy measures have an impact, in relative terms, on the supply of bank loans and induce changes in the financial structure of companies, which, in turn, affects their level of activity. As we explain in due course, the test does not, however, permit a distinction to be drawn between the two

Table 1

Structure of borrowed funds of non-financial firms (1987-96)

	Loans [1]	Foreign borrowing	Bonds	Short-term securities [2]	Trade credit	Other sources [3]
1987	47.8	7.2	7.5	2.7	21.0	13.8
1988	48.1	7.3	6.7	4.7	20.3	12.9
1989	49.6	5.8	6.1	3.9	20.5	14.1
1990	49.1	5.1	5.2	5.7	20.5	14.4
1991	49.5	5.7	5.1	4.4	20.3	14.8
1992	48.0	8.4	5.1	4.0	20.2	14.4
1993	45.3	11.4	5.4	3.3	19.4	15.2
1994	44.3	11.0	5.0	2.6	21.5	15.5
1995	44.3	10.4	4.4	2.4	23.6	15.0
1996	44.5	9.9	4.0	2.1	23.8	15.7

[1] Includes loans from national credit institutions. [2] Includes commercial paper and asset transfer certificates. [3] Includes credit from insurance companies, the public sector and households, technical insurance reserves and other deposits.

Source: Banco de España, Cuentas Financieras (1987-96).

meanings attached to the credit channel concept, although it does demonstrate the existence of a transmission channel which supplements the traditional money channel.

Vega (1992) reviews the role assigned to credit within the monetary transmission mechanism according to the economic and econometric literature. He provides estimates for the demand functions of credit and of ALP (liquid assets held by the public) in the Spanish economy, and he analyses the effectiveness of the credit controls introduced as a temporary measure between mid-1989 and early 1991. One of that study's findings is that, in connection with the controls, a process of substitution can be observed between bank loans and commercial paper. In this paper, our aim is to determine whether there are equivalent substitution processes in response to monetary policy measures – not necessarily quantitative restrictions. If so, this would provide evidence in favour of the existence of the credit channel.

Thus, two key questions to determine the existence of a bank lending channel impact are: (i) does the financing mix chosen by economic agents respond to changes in monetary policy, and (ii) does that mix affect their expenditure decisions?

The economic literature provides theoretical support and empirical evidence for a positive reply to the second question. For instance, non-fulfilment of the terms of the Modigliani-Miller theorem suggests that corporate financial structure does have a bearing on companies' investment decisions. The evidence in respect of the first question is less conclusive. The test discussed in this paper is an attempt to answer it. The article is structured as follows. Section 1 gives a brief description of the theoretical arguments in favour of the existence of the credit channel in the monetary policy transmission process. Section 2 provides, first, a brief review of the different approaches used to test the existence of that channel, and then presents the findings of the test based on the approach of Kashyap et al. (1993). It also assesses to what extent the findings are conditioned, in one part of the period studied, by the authorities' restrictions on the expansion of credit. The final section summarises the paper's conclusions.

1. Theoretical aspects

Several recent papers – Bernanke and Blinder (1988) and Kashyap et al. (1993), among others – offer theoretical models that support, in aggregate terms, the existence of a differentiated credit channel within the monetary transmission mechanism. This section, based on the more detailed studies carried out by Bernanke (1993), Gertler and Gilchrist (1993) and Kashyap and Stein (1994), summarises the arguments that constitute the theoretical basis for this transmission path.

The traditional view of the process through which monetary measures are transmitted maintains that reductions in the money supply induced by the authorities lead to an increase in real interest rates that is needed to restore equilibrium in financial markets. This increase in the cost of funding affects economic agents' expenditure decisions. As Cecchetti (1995) points out, this approach focuses on the impact on aggregate expenditure, but fails to take into consideration the possible differences in the effects that monetary policy has on different agents. By contrast, both versions of the credit channel approach address the problems of asymmetric information among economic agents with respect to financing, and emphasise the distributive consequences of monetary policy measures. Thus, the credit channel in the strict sense suggests that, for certain agents, the only way to surmount the problem of asymmetric information is to resort to financial intermediaries, i.e. specialised agents benefiting from economies of scale in monitoring tasks. Therefore, if monetary measures affect the supply of bank funds, some economic agents will be especially affected. Similarly, the credit channel in the broader sense indicates that asymmetric information problems lead to the existence of a risk premium determined by the value of the agents' net wealth. To the extent that monetary policy affects this net wealth, it will have a different impact on different agents.

1.1 The bank lending channel

The bank lending channel approach (or credit channel in the strict sense) stresses that monetary policy affects the level of economic activity not only by modifying short-term interest rates, but also by altering the availability and terms of bank loans.

In other words, whereas the money channel in monetary transmission refers to the effects on the liabilities of the credit system, i.e. the money supply, the credit approach emphasises what happens to the assets side of financial institutions' balance sheets. Hence, a monetary policy tightening, as it translates into a reduction in deposits, must be accompanied by a contraction on the other side of the banking institutions' balance sheets. This reduction in bank assets will have an effect on real economic activity additional to the impact of the money channel if two conditions are met:

(1) The monetary authority is able to affect the supply of intermediated credit. For this to be the case, there must be no other bank asset that could act as a perfect substitute for loans to companies.

(2) There exists no other alternative source of corporate financing that is a perfect substitute for bank lending.

If bank loans are an imperfect substitute for other bank assets, the counterpart to the drop in deposits will be a drop in the various kinds of bank asset (including loans) i.e. the banks will not be able to accommodate in full the agents' demand for funds solely by resorting to reductions in assets that are alternatives to credit. On the other hand, to the extent that firms (and consumers) lack perfect substitutes for bank loans, they will not be able to offset the lower availability (or different terms) of these loans simply by greater direct recourse to savers in public capital markets.

Both assumptions appear to be fairly reasonable. First, the different degrees of liquidity, profitability and risk of the various kinds of bank asset suggest that they are imperfect substitutes. Secondly, as a corporate liability, bank loans have no perfect substitutes, at least for a significant number of firms that depend heavily on bank financing and lack access to alternative sources.

Nonetheless, several authors – including Romer and Romer (1990), Thornton (1994) and Morris and Sellon (1995) – have cast doubt on the relevance of the two above-mentioned conditions, especially the first one. With regard to the first condition, factors such as the acquisition by banks of liabilities not subject to reserve requirements or the supply of credit from non-bank intermediaries would, in principle, weaken the possibility of a credit channel existing or, at any rate, its importance. Thus, Thornton and Morris and Sellon present evidence for the United States that suggests that financial innovation and deregulation processes have altered the structure of financial markets, weakening the monetary authorities' ability to control the supply of intermediated credit. Countering these arguments, Kashyap and Stein (1994) describe in detail several reasons why consideration of these factors is not so important. They underline, for instance, the fact that the marginal cost of financing via bank liabilities that are alternatives to deposits increases with the amount of financing obtained. They also stress that the volume of credit from non-bank intermediaries continues to be small compared with the volume of bank loans. As to the second condition, it tends to be weakened by companies' growing ability to tap non-intermediated funds or to obtain financing from abroad. Nevertheless, bank loans remain a principal source of funding for a significant number of firms.

1.2 The balance-sheet channel

Along with the previously described view of the credit channel (contraction of bank lending as a result of restrictive monetary measures), Gertler and Gilchrist (1993 and 1994) and Hubbard (1995) emphasise a related approach which, in looking at the impact of financial conditions on monetary policy transmission from another angle, generates several predictions similar to those derived from the existence of the bank lending channel. This complementary approach, usually known in the literature as the balance-sheet channel (or credit channel in the broader sense), rests on two basic ideas.

First, asymmetric information gives rise to a differential between the cost of internal financing and that of external funds. This external funding premium compensates lenders for the better information of the borrowers regarding the quality and profitability of investment projects. The second basic idea is that this differential between the cost of internal and external funds is inversely related to the net wealth that the borrower can provide as collateral. The greater the value of the collateral relative to the size of the loan, the greater the borrower's commitment to his own investment project.

Under these conditions, any disturbance affecting the net wealth of economic agents will affect the cost of external financing. Thus, an increase in the interest rate on open market transactions will lower the discounted value of assets that can be used as collateral, thereby raising the cost of borrowing. It will also increase the financial expenditure/cash flow ratio, thereby possibly reducing the volume of self-financing. Moreover, the initial drop in demand causes a subsequent decline in profits and the value of assets, with the resulting effect on the cost of external funds. Thus, this mechanism amplifies the impact of monetary policy measures. Unlike the bank lending channel, this credit channel in the broader sense does not depend on the institutional characteristics of the banking system. It is, rather, an operational transmission mechanism triggered by any disturbance affecting the net wealth of economic agents.

In sum, both approaches – the bank lending channel and the balance-sheet channel – can be understood as complementary mechanisms with which to explain, whenever asymmetric information problems exist, the influence of financial factors on the effects of monetary policy. Both would suggest a reinforcing of the traditional money channel in monetary policy transmission. But this reinforcement would take the form of a different impact on different agents, which would not be the case if only the money channel were operative. Thus, the credit channel in the broader sense tells us that there will be a greater impact on those agents for whom the cost of funds is more sensitive to the collateral offered, whereas the bank lending approach predicts a greater impact on agents who are more dependent on bank loans. Both approaches appear to suggest that consumers and small firms are the economic agents most affected.[1]

1.3 Credit rationing

Lastly, before beginning our analysis of the evidence on the existence of the credit channel, it is worth making one point about its relationship to credit rationing. The importance of credit in the monetary transmission process is not confined to situations in which credit rationing occurs. The existence of rationing has been well substantiated theoretically with models of asymmetric information in the credit market, and it is also an observable phenomenon, at least in certain periods. Nevertheless, credit rationing is not a necessary condition for the existence of the credit channel in either of its two versions. The bank lending channel will exist whenever conditions (1) and (2) are found to exist, i.e. that, faced with a restrictive measure, banks cut back loans and that firms are forced to supplement them by tapping alternative sources if they can, or by incurring additional costs (the increase in bank lending rates will exceed that of alternative instruments) if they continue to use bank financing. In other words, the bank lending channel is compatible with the existence of credit rationing, but it is also compatible with the increase in the relative cost of bank loans in response to monetary restrictions. At the same time, the balance-sheet channel will exist whenever the cost of borrowing is inversely related to collateral, even when there is no quantitative restriction.

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[1] In general, small firms rely more on bank loans, and they also bear higher borrowing costs. The latter may, as Gertler and Gilchrist (1993) point out, be due to the fact that they represent a higher risk since their activities are less diversified, to the relatively greater costs in the event of bankruptcy, and to the fact that, proportionally, they have less wealth to use as collateral.

2. Empirical evidence

2.1 Overview of the literature

Empirical studies of the credit channel have mainly been concerned with testing the existence of the bank lending channel in the monetary transmission process. To this end, a first approach, used in many studies, focuses on interpreting the correlations between money, credit and output, and on assessing how money and credit respond to monetary policy measures. For the US economy, the findings tend uniformly to show that credit responds to monetary shocks with a certain lag and that, in turn, changes in credit run fairly parallel to changes in output. Along these lines, Bernanke and Blinder (1992)[2] use a vector autoregression (VAR) model with six variables: the federal funds rate (FFR), the unemployment rate, the price level and three bank balance-sheet variables – deposits, loans and bond holdings. They find evidence that, after a monetary policy shock (measured as an orthogonal shift in the FFR), in the short run bond holdings adjust more strongly than loans, owing to the existence of committed loans (credit lines). In the longer term, the response of loans is greater and it tends to coincide in time with the response of the unemployment rate. Although this evidence is compatible with the existence of the credit channel, it is also compatible with the traditional money channel in the transmission of monetary impulses, which makes it difficult to identify which is which.[3] In other words, given a restrictive monetary policy, the observed drop in credit may be the result of a contraction in the supply of credit or it could be the result of a reduction in the demand for credit due to a slowdown in economic activity induced by increases in interest rates.

Another alternative found in the literature has been to assess the response of output to shifts in the volume of credit, either through studies based on the estimation of VAR models or by identifying periods in which there were clear shocks in the supply of credit (runs on the banks, for instance, or explicit credit controls) and measuring the response of output in such situations. The results of this approach show that contractions in the lending process induce significant fluctuations in output. This evidence thus proves that there is no perfect substitute – as a corporate liability – for bank loans; but in order to test the existence of the credit channel, it must also be proved that, as a bank asset, loans have no perfect substitute either.

A third approach, employed by Kashyap et al. (1993), attempts to solve the aforementioned identification problem by considering the relative behaviour of bank lending and some close substitute (in this case, commercial paper). Specifically, given a restrictive monetary policy, if the money channel were the only operative mechanism, one would expect the contraction in corporate demand for funds to lead to a simultaneous drop in all sources of finance. Conversely, if the credit channel in the strict sense also operates, the monetary contraction will induce a more pronounced decline in bank lending, which will tend to be substituted by alternative sources, at least by some of the companies with access to those sources.

The basic identification assumption on which this approach rests is that the contraction in the corporate demand for funds (induced by a contraction in aggregate demand) affects the different sources of financing proportionally. This assumption does not escape criticism. Thus, Eichenbaum (1994) suggests an alternative interpretation. Concretely, if small firms (which resort more to bank loans) were concentrated in sectors more sensitive to cyclical swings, they would face a greater fall in demand, and if they had more flexible technology enabling them to adapt production more swiftly to demand conditions, they would need to finance fewer stocks at the start of a slowdown than large

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2 Bacchetta and Ballabriga (1995) extend Bernanke's and Blinder's analysis to 14 OECD countries. For the majority of the countries, their findings are similar to those obtained by Bernanke and Blinder for the United States.

3 See Bernanke and Gertler (1995) for a critique of the comparative analysis of the dynamic response of monetary and credit aggregates to monetary policy measures, as an approach to testing the existence of the credit channel.

firms.[4] As a result, the demand for bank loans would decline in relative terms. More generally, as Kashyap and Stein (1995) point out, a relative increase in commercial paper could be observed, even if the credit channel were not operative, if the companies less affected by the contraction in aggregate demand are those that make greater use of this financing instrument.

Under this approach, an essential role is played by the variables related to the composition of corporate financial structure and, specifically, those variables that express the volume of bank loans in relative terms compared with their close substitutes. Employing a variable of this kind, Kashyap et al. (1993) carry out a two-stage test. First, an assessment is made of the response of the composition variable to changes in the degree of restrictiveness of monetary policy, with a view to examining compliance with condition (1) and using the above-mentioned identification assumption. If condition (1) is not found to exist, i.e. if all bank assets are perfect substitutes, then, faced with a contraction in their balance sheets, banks will be able to accommodate corporate demand for credit and the companies will not alter their share of liabilities. Conversely, if condition (1) holds, banks will not fully accommodate the demand for credit and companies will attempt to resort more to alternative sources of finance. Secondly, that paper examines whether the composition (mix) variable makes investment equations more meaningful, which is a way of testing condition (2). If this condition is not found to hold, i.e. if companies can replace bank loans with other sources at no extra cost, their financing mix will not affect their real decisions. Conversely, if condition (2) holds, financial structure variables can help explain investment behaviour.

As Oliner and Rudebusch (1995) point out, this test – specifically, its first stage – does not allow for a distinction between the broader and the strict notions of the credit channel. They maintain that the fall in the bank loans/alternative sources ratio may be due not just to a contraction in the supply of bank loans (as the strict version suggested), but also to the increase in the premium in the cost of all external financing (not just bank loans), especially for agents with greater problems of asymmetric information (as the broader version would predict). Thus, in their analysis of the reaction of the financial structure of small and large firms to monetary measures, they confirm that, after a restrictive monetary measure, there is a drop in the quotient between the volumes of indebtedness of small and large firms, while the bank loans/alternative sources ratio remains constant for both small and large companies. As small firms have a higher proportion of bank loans, the above findings imply a reduction of the bank loans/alternative sources ratio in the aggregate, even though there may have been no shift from bank loans to other sources in either small or large firms.

At the same time, tests for the credit channel in the broader sense discussed in the literature are based on the distinction between the behaviour of small and large firms when faced with monetary disturbances. Two examples of this kind of test are Gertler and Gilchrist (1994) and Oliner and Rudebusch (1996). The former analyse – for small and large firms – the different responses to restrictive monetary policy measures of a set of real variables (sales and stocks) and financial variables (short-term debt) and they observe that all are more pronounced in the case of small firms. Oliner and Rudebusch observe that investment expenditure relies more on self-financing following restrictive monetary shocks and that this reliance is significantly greater in the case of small firms. Nevertheless, it must be pointed out that the distinction by size does not allow distinguishing between the two versions of the credit channel, since, as discussed above, both versions would predict a greater impact on small firms: in the strict version because such companies are more dependent on bank loans, and in the broader version because their cost of financing is more sensitive to the collateral offered.

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4 A similar argument, discussed in Gertler and Gilchrist (1994), is that firms may play the role of marginal producers in some sectors. If this were the case, large firms would enjoy more stable demand and small firms would absorb residual demand and therefore be more sensitive to the cycle.

Finally, some recent papers make use of disaggregated data on bank balance sheets to implement a test of the bank lending view. These articles focus on the response of bank assets to monetary policy shocks, trying to assess the ability of the central bank to affect the supply of bank loans. The seminal reference in this literature is Kashyap and Stein (1995).[5] These authors find that the loan and security portfolios of large and small banks respond differentially to monetary policy shocks, which is consistent with the lending view.

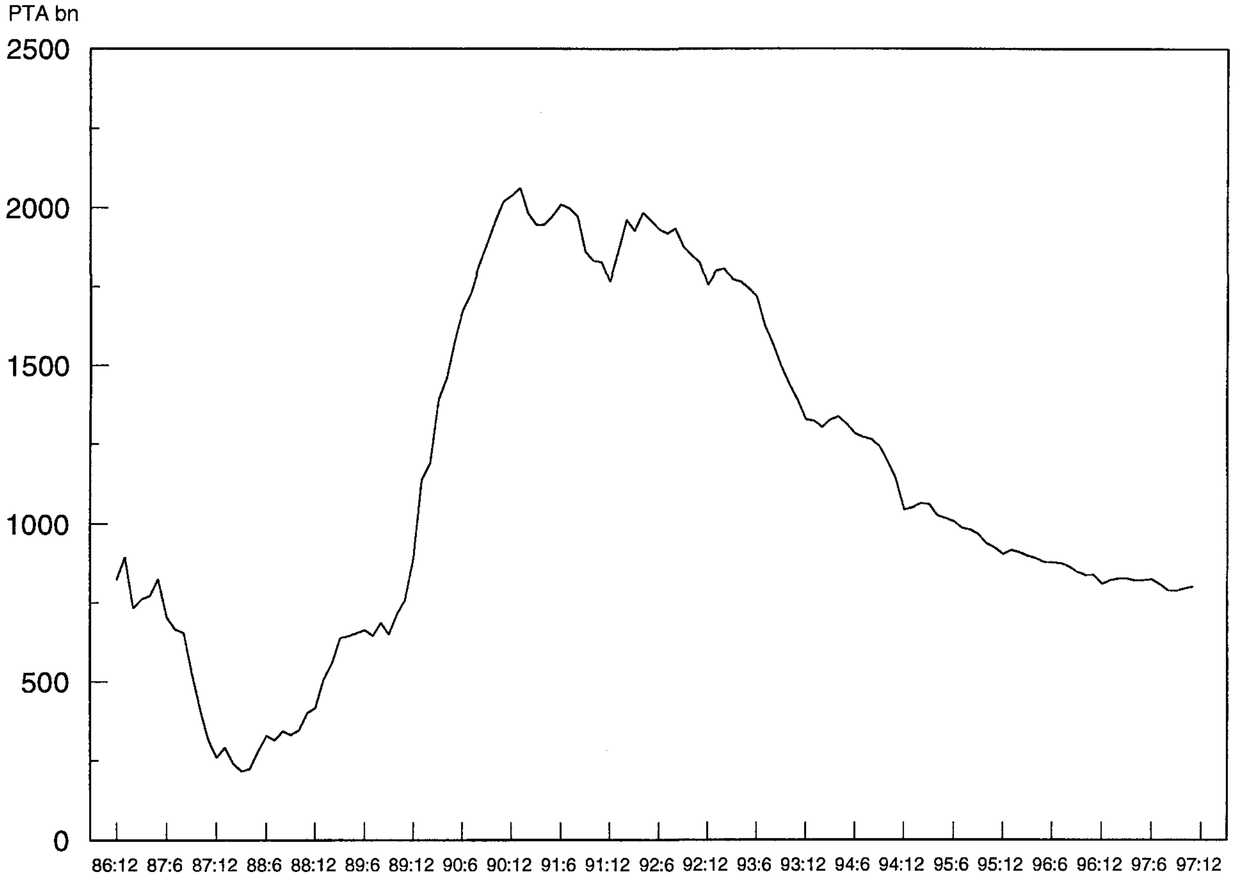
2.2 Monetary policy and corporate financial structure

In this section, we apply, for the Spanish case, the first stage of the Kashyap et al. (1993) test, which seeks to assess the validity of condition (1) by examining whether the response of bank loans to changes in monetary policy differs from that of some alternative asset. We forgo testing condition (2), because, as indicated in the introduction, the economic literature provides abundant evidence in favour of its fulfilment.

In the test put forward, the alternative asset chosen is commercial paper.[6] This is a short-

Chart 1

Commercial paper



Source: Banco de España.

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5 See Pill (1997) for an application of this approach to the Spanish case.

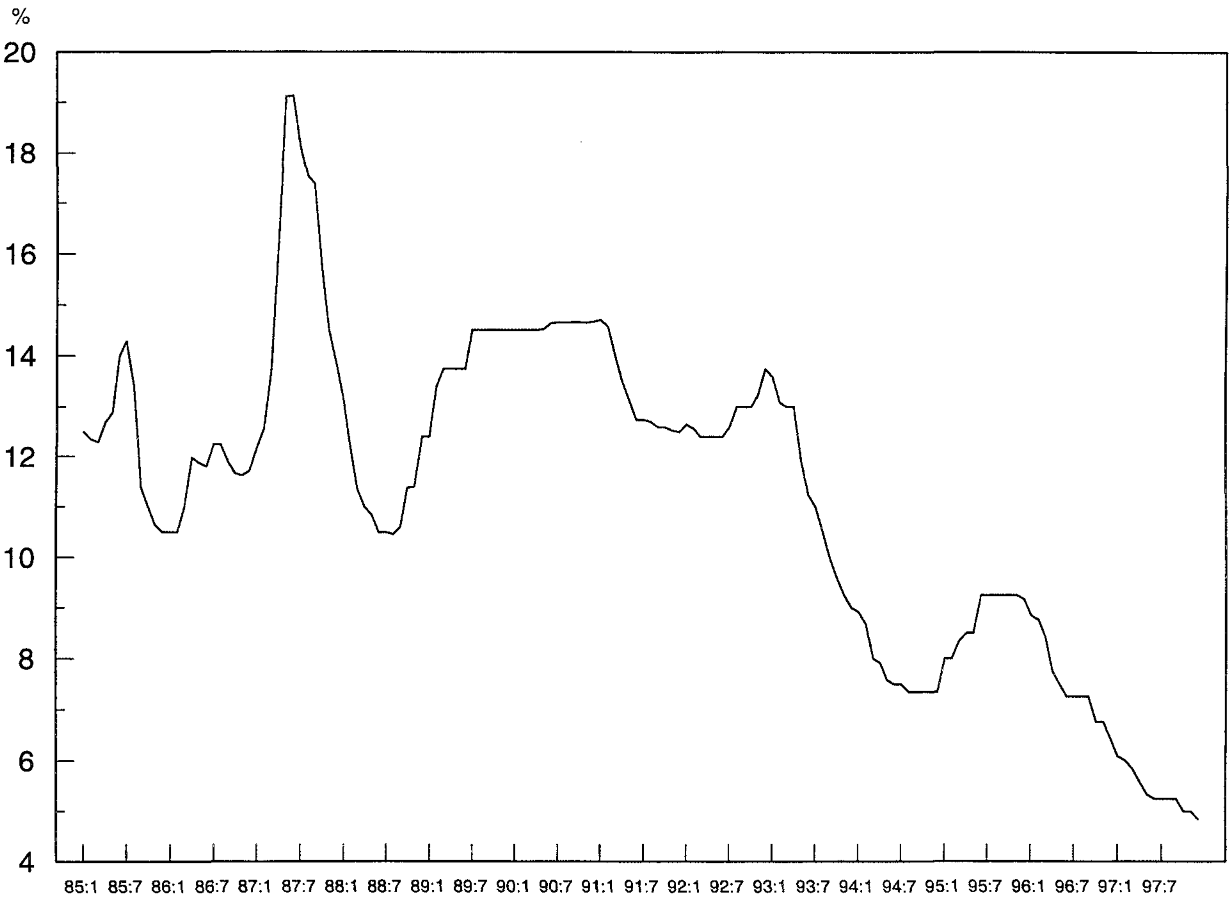
6 Other options were to use trade credit or foreign borrowing, which, according to Table 1, play a much greater role in corporate financing than commercial paper. The first alternative was discarded because only annual information covering a short period was available. The second option was disregarded owing to the significant changes in the regulations governing cross-border capital movements in the period under consideration.

term financing instrument recently created in the case of Spain and which has grown considerably in the past few years (see Chuliá (1992) for a detailed study of the Spanish commercial paper market). It has trended fairly unevenly owing to various disturbances. The 1985 law governing the tax status of certain financial assets made commercial paper, along with other assets, subject to tax withholdings. In 1987, the introduction of Treasury bills meant strong competition for commercial paper, especially given the high returns on the first issues of these bills. Lastly, in 1989, the introduction of restrictions on the growth of bank lending significantly boosted the market. The findings presented here were obtained using the monthly time series on commercial paper held by the public,[7] which is available from December 1986 (see Chart 1).

The choice of an indicator of the stance of monetary policy is somewhat more complicated. In some studies for the United States, the indicators used were discrete variables constructed after reviewing reports of the decision-making bodies of the monetary authority. Romer and Romer (1990) build a variable of this type, based on analysis of the minutes of the Federal Open Market Committee (FOMC), which identify the episodes in which there was a restrictive shift in the application of monetary policy. Another instance of this type of variable is the Boschen and Mills index used in Morgan (1993), which adopts five possible values to reflect the degree of tightness of monetary measures and which is also based on a review of FOMC minutes. Apart from these discrete

Chart 2

Intervention rate



Source: Banco de España.

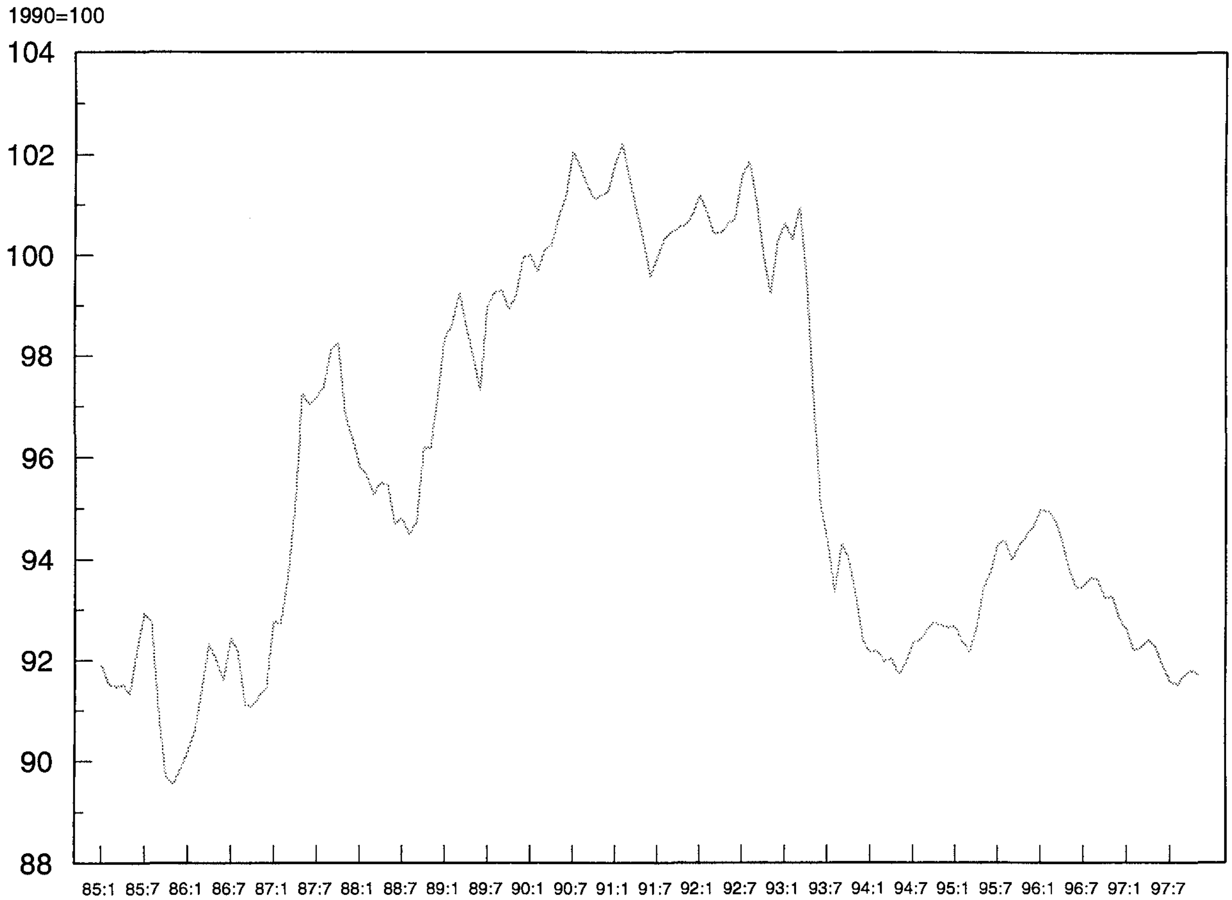
7 A description of the data used in the study can be found in the Appendix.

variables, in the case of the United States the level of changes in the FFR have often been taken as an approximation to reflect the stance of monetary policy since, with the exception of the 1979-82 period, the Federal Reserve has taken that rate as its reference in implementing monetary policy. Bernanke and Blinder (1992) and Balke and Emery (1994) produce evidence in favour of the suitability of the FFR as an indicator of the stance of monetary policy because, on the one hand, it is a good predictor of the main macroeconomic variables and, on the other, because it responds significantly to changes in the rate of inflation and the unemployment rate.

In the Spanish case, owing to the non-existence of a discrete variable, the choice is between a quantities variable and an interest rate. This paper, given the sample period used, opts for the Banco de España intervention rate[8] (see Chart 2). A further reason for this choice is the fact that, since the mid-1980s, the variable increasingly adopted as an instrument for monetary control has been a short-term interest rate rather than bank reserves, which had previously been closely monitored as a gauge for a given target (Ayuso and Escrivá (1993)). The intervention rate time series was taken as an indicator because using innovations in this rate – as Bernanke and Blinder (1992) do, and as Eichenbaum (1994) suggests as the appropriate way to identify exogenous monetary policy signals – requires calculating a monetary authority reaction function that is impossible to estimate given the sample period available. As a matter of robustness, a monetary conditions index (MCI, see Chart 3) is chosen as an alternative indicator of monetary policy.

Chart 3

Monetary conditions index

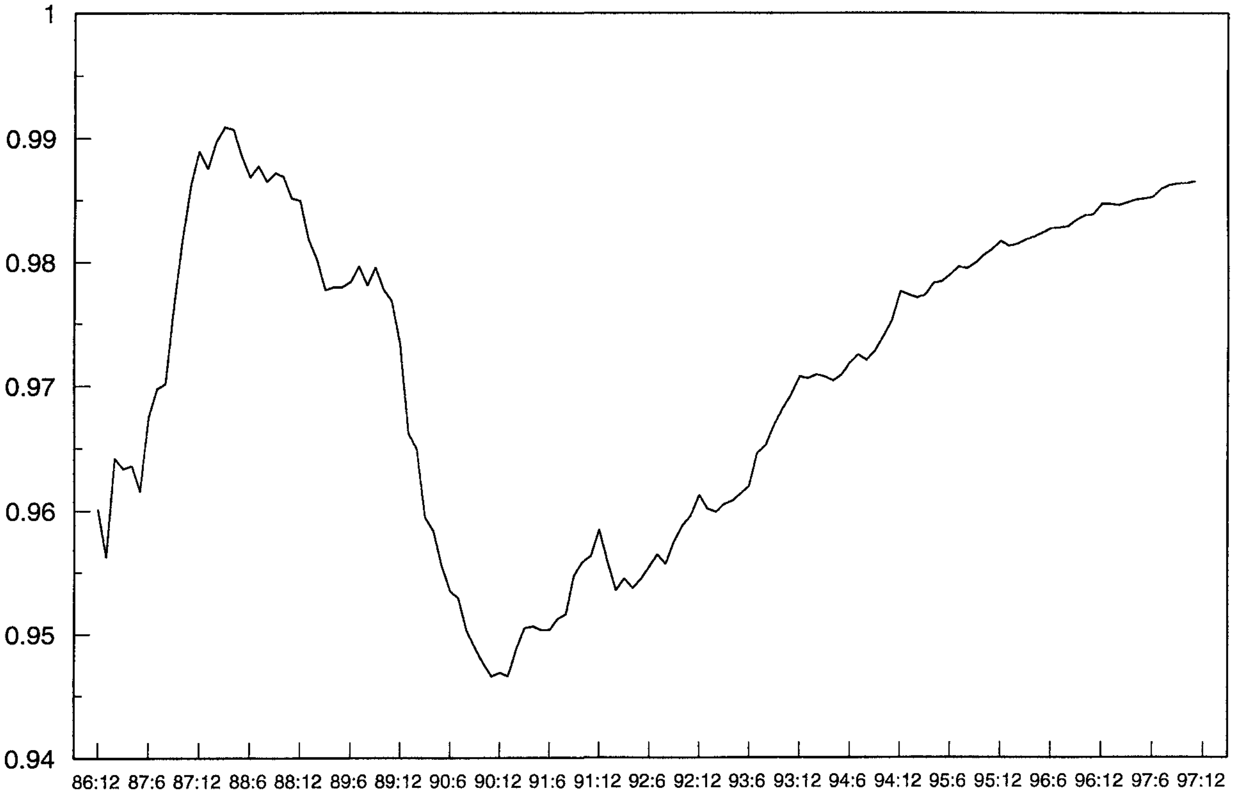


Source: Banco de España.

8 From May 1990 onwards, this is the rate applied for the ten-day repurchase tender for Banco de España certificates; to April 1990, it corresponds to the overnight assistance rate.

Chart 4

Composition variable



Note: The composition variable is defined as the ratio between bank loans and bank loans plus commercial paper.

Source: Banco de España.

In line with the first stage of the test presented in Kashyap et al. (1993), an attempt is made to assess the level of significance of the impact of the indicator of the stance of monetary policy on a composition variable i.e. the aim is to analyse whether monetary measures affect companies' choices of financing sources. The composition variable selected is the ratio between bank loans to residents (including firms and households) and bank loans plus commercial paper (see Chart 4). Table 2 sets forth the causality test findings. These are Granger-type tests and, to carry them out, the composition variable (or else the lending or commercial paper series) is regressed over 12 lags of its own and over 12 intervention rate (or, alternatively, MCI) lags. Alternatively, we have also included seasonal dummy variables or 12 lags of an activity indicator, specifically the industrial production index, in order to take into account seasonal or cyclical factors that could influence the composition variable. The test checks whether the lags in the monetary policy indicator help explain the behaviour of financial variables. Table 2 presents two types of statistic: first, statistic χ^2 (12) to test the null hypothesis that all the coefficients of the lags of the indicator of the stance of monetary policy are zero; and, secondly, the t-statistic for the sum of those coefficients. The first statistic represents the result of the Granger causality test, and the second reflects the sign and degree of significance of the accumulated effect of the proxy for the stance of monetary policy on the dependent variables.

The commercial paper series used to obtain the findings set out in Table 2 was the series for commercial paper held by the public. In implementing the tests, we use monthly data for a sample period ranging from 1989:1 to 1997:11.[9]

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[9] The available information allowed an estimation period starting in 1988:1 to be used. However, if the sample period begins in January 1988, by including lags in the regressors, we are using commercial paper data for 1987, a year in which

Table 2

Impact of the stance of monetary policy on corporate liability composition variables and on the relative cost of bank loans and commercial paper

$$\Delta x_t = \alpha_0 + \sum_{i=1}^{12} \alpha_i \Delta x_{t-i} + \sum_{i=1}^{12} \beta_i \Delta r_{t-i} + \sum_{i=1}^{12} \gamma_i \Delta \log A_{t-i} + seasonals$$

Dependent variable	***r* = intervention rate**		***r* = Monetary Conditions Index**	
	1	2	1	2
Commercial paper	54.99***	2.50**	30.78***	1.32
Bank loans	12.17	0.88	28.17***	2.24**
Composition variable [3]	47.28***	–2.33**	40.56***	–1.45
Spread [4]	59.74***	4.27***	44.35***	3.84***

Notes: x denotes the dependent variable (alternatively: commercial paper, bank loans, composition variable or spread), r the indicator of the stance of monetary policy (intervention rate or MCI), and A the industrial production index. Variables have been differenced so that they enter the regressions in stationary form. Specifically, the variables included are the logarithmic differences for commercial paper, bank loans (in real terms in both cases), and the industrial production index, and differences in the intervention rate, the monetary conditions index, the composition variable (MIX) and the loan/commercial paper spread. Sample period = 1989:1–1997:11.

[1] Statistic χ^2 (12) to test that all the coefficients of the intervention rate (or MCI) lags are zero (χ^2 (12) for $H_0 : \beta_i = 0 \ \forall_i$). Critical value at the 5% level = 21.03. [2] t-statistic for the sum of the coefficients of the intervention rate (or MCI) lags (t-statistic for $H_0 : \Sigma\, \beta_i = 0$). [3] The composition variable (MIX) is defined as the ratio between bank loans to resident sectors and bank loans plus commercial paper. [4] The spread is defined as the difference between the bank lending rate and the commercial paper rate.

The bank lending series includes, as mentioned earlier, credit to consumers and firms, for lack of monthly data on bank loans to companies. Given that the objective of the test is to analyse whether monetary measures affect the composition of corporate financing sources (and, specifically, the relative importance of bank loans), inclusion of credit to households could bias the results in favour of acceptance of the hypothesis, given the likelihood that consumer credit is more sensitive to monetary impulses. Conversely, the inclusion of long-term loans in the credit series used and the consideration of balances outstanding instead of new issues would tend to work against acceptance of the hypothesis to the extent that they imply a slower response by credit to monetary disturbances.

While the findings must be interpreted with the utmost caution, given the modest size of the sample and the aforementioned data limitations, it is worth pointing out some of the observations derived from Table 2. Starting with the tests using the intervention rate as an indicator of the stance of monetary policy, there is a clear rejection of the hypothesis that the coefficients of the intervention rate lags are zero in the commercial paper regression. What is more, judging by the t-statistic in the second column, there is a positive and statistically significant effect of the intervention rate on the behaviour of commercial paper.

As regards the regression with bank loans, the hypothesis that the intervention rate does not affect the volume of credit cannot be rejected, and, moreover, contrary to expectations, the sum of the coefficients is positive, although not significant. These findings may be explained, in part, by the evidence produced in other studies that credit responds with a certain time lag (sometimes longer than a year, which is the maximum lag considered in this paper) to monetary impulses.[10] This explanation is particularly relevant if, as in this study, the credit series includes long-term loans. On the other

the behaviour of commercial paper was clearly conditioned by the appearance of Treasury bills. As a matter of robustness, we repeated the tests for the period 1988:1–1997:11 by including a dummy variable of value 1 for 1988 (i.e. while the 1987 data appear as regressors), and the results did not change significantly.

[10] Thus Bernanke and Gertler (1995) point out that credit may rise in periods following a monetary policy tightening because consumers and firms may wish to soften the impact of cyclical swings on expenditure and output. Companies may, for instance, increase their volumes of stocks after a monetary tightening, going deeper into debt, albeit less than if credit markets were perfect.

hand, the findings are consistent with the fact that the intervention rate has increased in response to bouts of intense growth in lending in periods when economic activity has accelerated sharply.

The effects of the intervention rate on the composition variable are, because of the way the model is built, a combination of the previous findings. The hypothesis that the intervention rate does not affect the composition variable is rejected. At the same time, the sum of the coefficients for the intervention rate lags is, as expected, negative and significant. These findings suggest that monetary policy measures, by inducing changes in the conditions governing access to bank loans, lead to a recomposition of corporate liabilities.

The tests run with the MCI as indicator of the degree of tightness of monetary policy provide quite similar results. There is a clear rejection of the hypothesis that the coefficients of the MCI lags are zero in all the regressions. However, the sum of the coefficients of the MCI is now significant in the bank loans regression.

Complementing the above evidence are the findings of the test run on the causal relationship between the stance of monetary policy and the differential between the bank lending interest rate and the commercial paper interest rate. The hypothesis tested is that a monetary policy tightening leads to an increase in the relative cost of bank loans vis-à-vis commercial paper. A finding in favour of this hypothesis would be consistent with the evidence derived from the tests with the composition variable, since we would expect to find a close inverse relationship between the loan/commercial paper spread and the composition variable defined above.

The last row in Table 2 sets out the findings of this test aimed at ascertaining whether intervention rate lags (or, alternatively, MCI lags) help explain the behaviour of the bank loans/commercial paper spread. This spread, with and without adjustment for fiscal factors, is depicted in Chart 5.[11] There are two basic findings:[12] first, the hypothesis that the coefficients of the intervention rate are zero is rejected; and, secondly, there is a positive effect of the intervention rate on the spread considered.[13] These results hold when the MCI is used instead of the intervention rate. This evidence might be partially explained by the small size of the Spanish commercial paper market, which may induce a sluggish response of commercial paper rates to monetary policy shocks.

The results shown in Table 2 appear to point in the expected direction in the sense that, in response to a monetary policy tightening (identified by increases in the intervention rate or in a monetary conditions index), there is a relative increase in the cost of bank loans vis-à-vis the cost of commercial paper, and, in terms of the financing sources, a drop in the relative proportion of bank loans can be observed. These findings, in line with those obtained in Kashyap et al. (1993), thus argue in favour of the hypothesis that an active credit channel exists in the transmission of monetary impulses.

Given the time period taken, which includes a prolonged phase (relative to the sample size) of government restrictions on the growth of credit, it is essential to evaluate the extent to which the results obtained are conditioned by the existence of this period of direct controls on credit growth. To this end, the previous analysis was repeated by incorporating in the regressions with the intervention rate a dummy variable which assigns the value 1 for those months during which quantitative restrictions on credit were in effect. As expected, this dummy variable has a positive and significant effect on commercial paper and a negative effect on bank loans and the mix variable,

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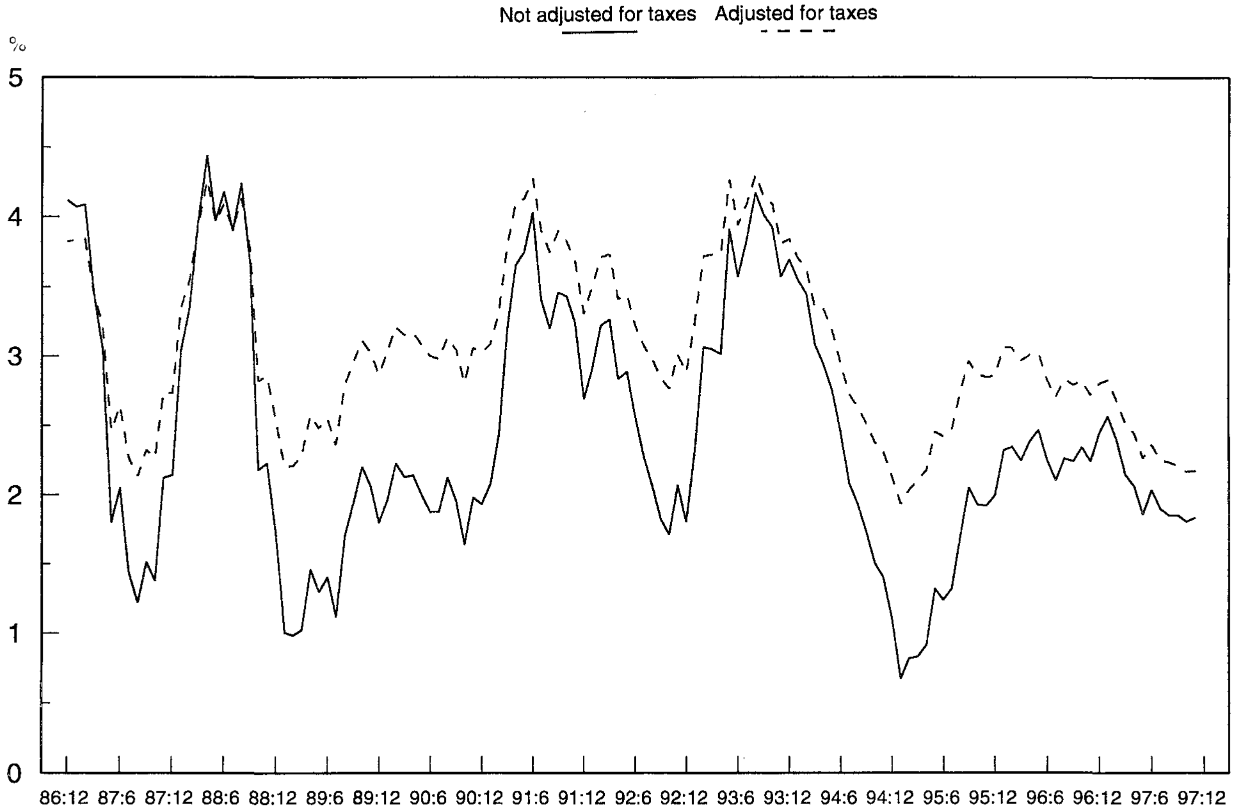
11 Details of the adjustment for fiscal factors applied to bank lending and commercial paper rates are given in Cuenca (1994).

12 The results of the tests with both definitions of the spread are very similar. In Table 2, the results with the spread without tax adjustment are presented.

13 This finding is consistent with the increasing sensitivity of the bank lending rate to changes in the intervention rate since the late 1980s (see Box V.3 in Banco de España (1994)).

Chart 5

Domestic loans/commercial paper spread



Source: Banco de España.

although in the case of loans the effect is not significant. At the same time, that dummy variable has a negative, although not significant, effect on the differential between the lending rate and the rate for commercial paper. Two possible ways to explain that sign are, first, the intense demand pressure in the commercial paper market during the period of credit restrictions; and, secondly, the possibility that, under those circumstances, banks might have considered that increases in their lending rates would translate into notable increases in the average risk of the projects they were financing.

The results of the causality tests – which took into account the phase of direct controls on credit growth – are shown in Table 3. With regard to the quantity variables, Table 3 shows, in the commercial paper and composition variable regressions, that the hypothesis that the coefficients of the intervention rate lags are zero continues to be rejected. The effect of the intervention rate on commercial paper continues to be positive, albeit less significant. Similarly, the impact on the composition variable is negative, but ceases to be significant. As to the relative cost of bank loans and commercial paper, a comparison of Tables 2 and 3 shows that the results do not change substantially and maintain the positive and significant effect of the intervention rate lags on the loan/commercial paper spread.

Comparing these results with those shown in Table 2, two conclusions can be drawn. First, the observed decline in the relative proportion of bank loans may be more pronounced in response to the introduction of direct controls on the growth of credit than in response to increases in the intervention rate. Secondly, the relative increase in the cost of bank loans vis-à-vis commercial paper in response to increases in the intervention rate appears to be a solid finding which also holds when the period of direct credit controls is taken into account.

Table 3

Impact of the stance of monetary policy on corporate liability composition variables and on the relative cost of bank loans and commercial paper

$$\Delta x_t = \alpha_0 + \sum_{i=1}^{12} \alpha_i \Delta x_{t-i} + \sum_{i=1}^{12} \beta_i \Delta r_{t-i} + \sum_{i=1}^{12} \gamma_i \Delta \log A_{t-i} + seasonals + \psi_0 DUMCR$$

Dependent variable	***r* = intervention rate**	
	1	2
Commercial paper	50.53***	2.12**
Bank loans	12.22	0.87
Composition variable [3]	50.00***	–1.04
Spread [4]	59.53***	4.32***

See notes to Table 2. The difference with respect to Table 2 is that the regressions on which the above results are based incorporate as a regressor a dummy variable (*DUMCR*) that takes values of 1 for observations in the July 1989 to December 1990 period.

The evidence so far presented is consistent with the findings in van Ees et al. (1994) for the Dutch case, which underline the lag in the response of bank loans to indirect monetary control measures and, hence, the lower degree of effectiveness of the credit channel in this case as compared with the introduction of direct controls. This finding would help explain the first of the conclusions discussed in the previous paragraph, especially if we bear in mind that we are considering the effect of the intervention rate on bank loans with a maximum lag of one year. Regarding the relationship between the credit channel and alternative ways of implementing monetary policy, it is worth looking at Tsatsaronis (1994), who analyses, for the past three decades, the existence of the credit channel in four countries: Germany, the United States, Japan and the United Kingdom. A priori, the credit channel might be expected to be more important in Japan and Germany, countries characterised by a high degree of bank involvement in corporate projects and where bank lending plays a greater role in corporate financing than it does in the United States or the United Kingdom. Nevertheless, Tsatsaronis finds evidence that the credit channel is most effective in the cases of Japan and the United Kingdom, both of which resorted, especially prior to 1980, to direct credit control mechanisms. This finding suggests that, as far as the effectiveness of the credit channel is concerned, more importance attaches to the type of monetary policy instrument used than to the relative weight of bank lending.

Conclusions

The analysis of the existence of a credit channel in the transmission of monetary impulses is a frequent topic in recent economic literature. Nevertheless, this area of research continues to suffer from certain limitations. On the theoretical level, there is a marked lack of microeconomic substantiation and a lack of precision in defining the credit channel concept. At the same time, the various empirical approaches, using alternative methodologies, are inevitably based on identification assumptions that are needed to determine to what extent changes in credit respond to alterations in credit supply or demand. Another limitation lies in the inability of the available tests to distinguish between the different transmission mechanisms that are classified together under the concept of credit channel. Finally, although there is some agreement regarding the existence of the credit channel, the quantification of its impact and the assessment of its importance in relation to the liquidity channel have yet to be addressed with due rigour.

In addition to briefly reviewing the theoretical arguments associated with the existence of the credit channel and distinguishing between the different meanings attached to it, this study conducts a test of its existence based on Kashyap et al. (1993). Specifically, it tests the non-existence of alternative assets that could be perfect substitutes for bank lending as a bank asset, in such a way that restrictive monetary policy measures manage to reduce the supply of bank loans, and financial institutions do not fully accommodate the corporate demand for funds. If this condition is met and if companies do not have a perfect substitute for bank loans, the credit channel will operate.

More precisely, the findings of this study show that, in the event of a monetary policy tightening (for which the Banco de España intervention rate and, alternatively, a monetary conditions index serve as proxies), there is an increase in financing via commercial paper relative to funds obtained from financial institutions. Consistent with this, from the standpoint of prices, is the fact that a relative increase in the cost of bank loans vis-à-vis the cost of commercial paper is observed.

When the existence of a period of direct controls on the growth of credit is taken into account and its effect is excluded, the impact of changes in the intervention rate on the differential between lending and commercial paper rates is maintained, but its influence on the composition of corporate liabilities is more moderate. This finding partially weakens the evidence for the existence of the credit channel.

These findings are conditioned by the short sample period used and by the definition of credit employed (which includes lending to households and long-term loans). As indicated, the shortness of the sample period is especially worrying, given the possibly excessive weight in the analysis of the period of government restrictions on the growth of bank lending.

In addition, the short sample period available means that the analysis cannot consider effects with a time lag of more than one year. Given that the full response of lending to monetary policy measures can take longer, consideration of a larger number of lags would argue in favour of the findings obtained, i.e. acceptance of the hypothesis that changes in the stance of monetary policy affect corporate financial structure.

Finally, the tests presented in this paper cannot discriminate between the different versions of the credit channel concept. The use of microeconomic data is essential to distinguish between these alternative versions.

Appendix: data used

- Commercial paper held by the public. *Source*: Banco de España, Boletín Estadístico.
- Bank loans: we used the time series on the loans of credit institutions to other resident sectors (private sector). *Source*: Banco de España, Boletín Estadístico.
- Composition variable: defined as the ratio of bank loans to bank loans plus commercial paper.
- Banco de España intervention rate: to April 1990 we used the overnight assistance rate, and from May 1990 onwards the rate for the ten-day repurchase tender for Banco de España certificates. *Source*: Banco de España.
- Monetary conditions index: constructed with the weights being 1.5 to 1, interest rate to exchange rate.
- Industrial production index: general index with base year 1990. *Source*: Instituto Nacional de Estadística.
- Bank lending rate: we used the time series on domestic credit extended by banks and savings banks. *Source*: Banco de España.
- Commercial paper rate. *Source*: Banco de España.

On the method used to construct these time series on interest rates, with and without adjustment for fiscal factors, see Cuenca (1994).

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Comments on: "The credit channel in the transmission of monetary policy: the case of Spain" by Ignacio Hernando

by K. Alec Chrystal*

It is a great pleasure to have been asked to read and comment on this paper which addresses an important and interesting topic. It is also one of great importance for monetary authorities around the world as it potentially helps the understanding of the transmission mechanism of monetary policy. The traditional view has the transmission mechanism working via interest rate and exchange rate effects, which then influence components of aggregate demand such as investment or net exports. The credit channel view adds another influence via the supply response of the banking system. In other words, there is a quantity effect that is an additional determinant of spending and looking at market prices alone does not capture this effect. In essence, it is argued, banks change their policy on loan supply at various times and this means that spending will be lower than it otherwise would be *at a given level of interest rates.*

The paper falls into two parts. The first part provides a survey of the credit channel literature and the second part has some empirical evidence for Spain. I have nothing more to say about the first part, as this is an excellent review of the literature, except to recommend it to all readers. On the second part, I am obliged to offer some comments and criticisms as this is the job assigned to any discussant.

The empirical work is based upon a study of the change in composition of aggregate lending between bank lending and commercial paper. The point to be established is that a change in monetary policy affects the composition of finance, with the implication that changes in the *relative* role of banks versus securities markets indicate the existence of a credit channel. I will discuss the specific evidence offered first, and then think aloud about whether this kind of evidence could ever help to tell us much about the existence or not of a credit channel.

The core of the empirical work uses Granger causality tests to determine whether bank lending, commercial paper (CP) and CP as a proportion of total lending are caused by changes in the monetary policy intervention rate. Also tested is the Granger causality of the policy intervention rate on the spread between the interest rate on bank loans and commercial paper.

The strong conclusion from the empirical work is that the composition of lending is significantly caused by the intervention rate, as is the spread between the bank lending rate and the CP rate. This seems to be a robust result, as the statistical significance levels are very high. However, the question we need to answer is: what do we make of this? And in particular: what does this tell us about the credit channel?

There is considerable doubt about whether we can draw any conclusion at all from the results relating to the composition of lending. One problem is that the CP market is very small relative to the bank loan market, varying between 1 and 5% of the total. More importantly, perhaps, the swings in the CP market that did occur are mainly explained by institutional factors, such as the 1985 tax law, competition from Treasury bills in 1987, and bank lending restrictions in 1989 (as explained by Hernando and illustrated in his Charts 1 and 4). This makes it very hard to believe that changes in the composition of borrowing and changes in the level of CP issuance are primarily related to changes

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in the intervention rate (a variable that moves with a very different pattern, as shown in Hernando's Chart 2).

This does not mean that the intervention rate is not a significant influence on the composition of lending, merely that we would need a longer series of data for periods not subject to the institutional factors mentioned above before we could be convinced. If forced to reserve judgement on the influence of the intervention rate on the composition of lending, this leaves the affect of the intervention rate on the spread as the main robust statistical finding. So, does the fact that the spread is clearly positively correlated with the intervention rate prove the existence of a credit channel?

It is easy to remain sceptical. This does not mean that it is not consistent with a credit channel, but merely that there may be other interpretations.

At one level, this evidence supports the credit channel literature as it is clear that *all* bank loans and CP cannot be perfect substitutes. Indeed, we probably did not need evidence of this fact because if all bank loans and CP were perfect substitutes there would be no reason for banks to exist in the first place. Disintermediation would occur and savers would lend to firms via securities markets. Of course, banks do exist and the reasons are well known. Most importantly, securities markets are only a potential source of loans for large well-known companies where counterparty risk is minimal. Banks provide finance mainly for persons and small businesses *who cannot access securities markets.*

This causes a problem, because, in comparing the interest rate on CP with the average interest rate on all bank loans we are not comparing like with like. This is because (I presume) the interest rate quoted for bank loans is an average rate and is not necessarily the rate that banks would quote for loans to the very top corporate clients. If we could identify such a rate we might find that it moved very closely with the CP rate. Indeed, if it did not, why would corporate treasurers borrow at a higher rate than they needed to. Certainly, in the United Kingdom, large corporates can borrow at rates very close to the rates that banks themselves would pay in the interbank market. If this is true, then changes in the spread between average bank loan rates and CP rates are not an indicator of any quantitative change in the supply of loans offered by banks to top corporates. Rather, changes in the spread are an indication of pricing of bank loans to customers who have no alternative source of finance.

In order to draw strong conclusions from the behaviour of bank spreads, it seems to me that we need more detail on the behaviour of banks in the market for their liabilities. What, for example, happens to bank deposit rates when the intervention rate changes? Does this spread widen, and if so what does that tell us about the relative competitiveness of the markets for retail deposits and loans?

It could be, for example, that bank spreads (over intervention rates) rise when monetary policy is tight because the default risk on loans goes up. If this were the case, it would be a response that is not at all inconsistent with an interest rate transmission mechanism so long as countercyclical credit spreads are built into the story. Certainly, there is no unambiguous support here for a credit channel story, as there is a plausible "equilibrium" pricing story that is consistent with the same data.

Another way of making the same point is to note that banks have some discretion in the choice of loan rates they set. Observations on how this loan rate moves relative to money market rates over the cycle do not confirm or deny the existence of some form of quantity rationing that is required for the credit channel to exist. It may still be there, but, then, it may not. We still have the classic identification problem that has always haunted the search for a credit channel when using aggregate data.

None of this should be taken as an expression of disbelief that there is a credit channel. On the contrary, I take it to be so self evident that it has existed and could exist from time to time that I find it strange that such a large literature has focused on the not very interesting question of existence. Credit crunches clearly do happen. There was probably a credit crunch in the United States in the early 1990s and there was one in the United Kingdom in 1974/5 at the time of the secondary banking crisis. Equally, there was a credit surge in the United Kingdom when the corset was removed in 1980. At such times there are shifts in the bank loan market that are not captured by interest rate changes alone. Clearly, we need to understand such forces at times when there are major institutional or regulatory reforms.

The real question, however, aside from periods of structural change or banking crises, is: what else to we need to tie down about the credit channel that can help us understand the transmission of monetary polcy in normal times? If credit channel effects move in a sytematic way in response to movements in policy-determined interest rates, then we have nothing to worry about because the credit channel is just part of the econometric linkages from interest rates to activity that we have been trying to estimate all along.

One line of attack on this issue that is being followed in the Bank of England is to study the determinants of bank lending at a sectoral level, and to model these simultaneously with specific aggregate expenditure functions. This can potentially tell us if there is information in bank lending data relevant to explaining private sector spending that we have not been picking up through more traditional approaches (such as interest elasticities of expenditures, and money demand studies).

In conclusion, it is clear that Hernando has taken on an extremely difficult task in trying to identify credit channel effects in aggregate data. This is difficult enough where there are long runs of data. It is even harder in the case of Spain, where there have been structural changes that make hypthesis testing difficult. These difficulties (and the other problems mentioned above) should not be taken as a discouragement. On the contrary, work of this kind is extremely valuable. I have certainly learned a great deal by reading it. I hope that the author continues with work of this kind.

Monetary policy, aggregate demand, and the lending behaviour of bank groups in Switzerland

Olivier Steudler and Mathias Zurlinden*

Introduction

According to the conventional interest-rate transmission mechanism presented in most textbooks, monetary policy operates through changes in the interest rate and the credit creation process can be ignored. In contrast, the bank lending view argues that monetary policy by changing the volume of reserves influences banks' ability to lend and through this the real spending of bank borrowers. The two channels are not exclusive. Most economists and observers of monetary policy accept the existence of the interest-rate channel, and disagreement is usually confined to the question whether the bank lending channel really matters. The answer has consequences for monetary policy. If the bank lending channel works central bank actions may influence aggregate demand even in the absence of a change in the interest rate.

The empirical literature on the bank lending channel generally has focused on the correlations between monetary policy variables, aggregate demand, and bank loans. The evidence is not conclusive however. In particular, there is the problem of disentangling loan supply effects from loan demand effects. Dissatisfaction with this situation has prompted some economists to experiment with disaggregated data. This paper follows Kashyap and Stein (1995), who looked at the differential effects of monetary policy on the investment policies of small versus large banks with US data. Kashyap and Stein argued that large banks have better access to non-deposit forms of external finance. As a consequence, large banks will reduce their loans slower than small banks and are more inclined to use their securities holdings as a buffer during a monetary contraction. In this paper, we test this prediction for Switzerland.

The strategy is to estimate small unrestricted VARs and to examine the impulse response functions. In particular, we investigate the responses of cross-sectional differences in bank lending behaviour to innovations in interest rates and in bank deposits. The three groups of Swiss banks considered are *Big banks*, *cantonal banks*, and *regional banks*. These groups are used as proxies for banks of different size. The balance sheet data are quarterly and refer to claims and liabilities on residents in Swiss francs. The estimation period is 1977:2-1996:3.

By confining ourselves to the traditional classification of the Swiss banking sector, we may omit potential information. In return, we obtain a longer time series, which covers more than one business cycle. Complete balance sheet data for individual banks are available for the period since February 1987 only, while aggregated series for the three bank groups start in the 1970s. We use the original aggregated series of this traditional classification for the period 1975-87 and reconstruct the time series from scratch for the 1987-96 period. This procedure is sensible because many banks have merged or have been taken over by other banks during this second period.

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1. Background

In recent years, there has been considerable interest in the credit view of the transmission mechanism. Useful discussions of the theoretical and empirical literature are provided by Cecchetti (1995), Hubbard (1995), and Kashyap and Stein (1995). Somewhat simplified, we can distinguish between two variants: the balance-sheet channel and the bank lending channel. It is only in the bank lending channel that banks and bank credit play a special role. Yet the literature on the balance-sheet channel has influenced the way the bank lending channel is examined in this paper and is therefore discussed first.

The balance-sheet channel asserts that information problems between the borrower and the lender drive a wedge between the price of uncollateralised external funds and the price of internal funds. This premium is inversely related to the net-worth of the borrower. A contraction of monetary policy, reflected in a rise in short-term interest rates, increases the debt burden of the borrower and reduces the value of his collateral. The premium that must be paid for external finance rises and borrowers will curtail their real spending. Thus, monetary policy, by affecting the balance sheet of the borrower, can influence aggregate demand.

The second variant of the credit view stresses that monetary policy can have an impact on the supply for bank loans. Bernanke and Blinder's (1988) popular model of the bank lending channel is an IS-LM model with three assets, i.e., money, bonds and loans. The resulting picture of the transmission mechanism is as follows. A contraction of monetary policy by reducing reserves forces the banks to decrease their deposits. Lower deposits, in turn, trigger an adjustment of the asset side. Since bonds and loans are imperfect substitutes, the banks will attempt to reduce both forms of assets. Firms and households face a smaller supply of loans and have to reduce their investment projects.[1]

Notice that imperfect substitutability must hold for banks as well as for firms and households. This implies that at least some firms or households must depend on bank loans as a source of external finance. They cannot get hold of funds from other sources without additional costs. Equally, banks cannot compensate their loss of deposits by funds from other sources without additional costs. They cannot completely isolate their loan portfolio.[2]

In an empirical paper, Bernanke and Blinder (1992) examined the effects of US monetary policy on deposits, securities, and loans of the US banking sector. They estimate an unrestricted VAR and find that the deposits and securities fall immediately after a rise of the federal funds rate, while loans and output decline with a considerable lag. These results have been corroborated by McMillin (1996) for the United States, and Bacchetta and Ballabriga (1995) for a sample with 14 European countries.

Kashyap and Stein (1995) applied the same strategy on disaggregated US data. They construct bank groups by size and look at how deposits, securities and loans of these bank groups respond to monetary policy shocks. The basic assumptions are that the financing with non-deposit forms of external funds has rising marginal costs and that these costs are larger for small banks than for large banks.[3] Kashyap and Stein derive two predictions for a homogeneous competitive loan market. First, after a reduction of deposits loans of small banks decline more rapidly than those of large banks. Second, the securities of small banks decline less rapidly than those of large banks after a reduction of deposits. This implies that large banks are more willing than small banks to use their securities as a buffer when confronted with a monetary policy shock. Small banks value their

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1 For extensions of the Bernanke-Blinder model, see Keeton (1992).

2 Romer and Romer (1995) have doubted the validity of the second assumption.

3 It could be argued for example that small banks face higher agency costs. See Myers and Majluf (1984), and Stein (1995).

securities higher because they face higher marginal costs of external funds. As a consequence, they are more willing than large banks to adjust their loans portfolio. The results presented by Kashyap and Stein (1995) are consistent with this story for loans; the results for securities are mixed.

The main motive for the disaggregation of the banking sector is to disentangle shifts in loan demand from shifts in loan supply. Gertler and Gilchrist (1994) pursued a similar strategy in a study on the response of small versus large manufacturing firms to US monetary policy. Kashyap and Stein's basic insight was that there is no fundamental difference between the bank's access to non-deposit forms of external finance and the firm's access to uncollateralised external funds. Disaggregation does not completely eliminate the identification problem, however. As Kashyap and Stein readily concede, it could still be argued that large banks lend to large customers whose loan demand is less cyclical. As a result, large banks should have smaller swings in loans than small banks because of heterogeneous demand. Kashyap and Stein (1997) tackle this issue in a subsequent paper by holding bank size fixed and by focusing on the differences in balance-sheet strength within a given size category. Their study is based on a panel data set that includes quarterly observations of every bank in the United States over the period 1976-93. Such a study is beyond the scope of this paper.

2. The bank balance sheets and the construction of the data

The construction of the data used in this study was a laborious task and is described here in some detail. The balance sheets of the various bank groups are then characterised, and the question of whether our sample is representative for the Swiss banking sector is briefly discussed. Information on the evolution of the resulting time series over time is given in the next section.

The data are based on the monthly balance sheets of banks in Switzerland and Liechtenstein with total assets of more than 150 millions Swiss francs (100 millions until 1993). These balance sheet data are available for each individual bank for the period since February 1987. For earlier periods monthly data exist for bank groups only (Big banks, cantonal banks, and regional banks). Useable aggregate figures for these traditional classification are available since July 1975 when the structure with subdivisions into "residents" and "non-residents" and "Swiss franc" and "foreign currencies" was adopted. The three bank groups were defined in the early 1970s, resulting at the time in 5 Big banks, 28 cantonal banks and 39 regional banks. Since 1975, the number of banks included in the sample has decreased, mainly because of mergers and take-overs between banks. Banks that were dropped from the sample were not replaced by newcomers. Two points are noteworthy. First, the concentration process led to a sharp decline in the number of regional banks, and an increasing dominance of the Big banks. Second, the share of the sample in the total has not changed much over time. The banks included in the sample were 72 in 1975, representing 77% of the Swiss banking sector (measured by total assets). At the end of 1996 the corresponding figures were 47 and 78%.[4]

With respect to mergers and takeovers, we make use of the observation that the number of banks included in the sample declined by one between July 1975 and February 1987, and by 24 between February 1987 and the end of 1996. Thus, the balance sheets of individual banks can be used to construct time series adjusted for the effects of mergers and takeovers for the period since February 1987. This is done because we are interested in the behaviour of banks over the business cycle. By taking out the jumps caused by the concentration process we tacitly assume that mergers and takeovers do not reflect the effects of monetary policy.

The construction of the time series is based on a procedure proposed by Kashyap and Stein (1995) and adapted for our purpose. For every two adjacent months, those banks among the 71

[4] A table with the banks included in the sample can be provided on request.

banks of our gross sample are identified, which are not involved in a merger or take-over in the second month, and for which complete data are available for both months. Then, the banks are assigned to the three groups, aggregated balance sheets are constructed for each group and both periods, and rates of change are calculated for each of the three groups and for all balance sheet items we are interested in. The same procedure is then repeated for the next pair of periods. The procedure is applied for all periods from February 1987 to November 1996.

Table 1 provides some basic information on the three bank groups' balance sheets. The date is February 1987; i.e., roughly the mid-point of the sample and the first date balance sheets for each bank are available. Several patterns emerge from Table 1. Big banks have relatively large claims and liabilities vis-à-vis non-residents. They also have relatively large assets and liabilities in foreign currencies. Both components are of little importance for cantonal banks and regional banks. For these two bank groups it does not make much difference whether the total balance sheet or the domestic Swiss franc component is focused on.

On the asset side Big banks hold a relatively large share as interbank time deposits and securities (mainly in foreign currencies), while the asset side of cantonal banks and regional banks is dominated by secured loans (mainly mortgages). If we focus on the Swiss franc claims on residents, the Big banks still hold a relatively small share of their assets as secured loans. On the other hand unsecured loans are more important for Big banks than for cantonal banks and regional banks. This is true for the overall total and for the Swiss franc claims on residents.

On the liability side the Big banks owe relatively large shares as sight deposits or time deposits (again mainly in foreign currencies). Cantonal banks and regional banks finance their business mainly through bonds and saving deposits. The large share of bonds may come as a surprise and contrasts with the data for the United States reported by Kashyap and Stein (1995). We examined this issue by ordering the balance sheets of all Swiss banks with total assets of more than 100 millions Swiss francs in accordance with size and by constructing asset size groupings for large banks (in the 98th percentile) and various degrees of smaller banks (defined as those at or below the 95th, 90th, and 75th percentile). The resulting figures (not included in the paper) indicate that the Big banks have particularly low bond obligations. The other banks line up as predicted with small banks having the lowest share of bonds.

Finally, a comparison between the balance sheets of the traditional groupings and the asset size groupings may help answer the question whether Big banks, cantonal banks and regional banks are representative for large, medium, and small banks, respectively. The results indicate that the Big banks coincide to a large degree with the class of the largest banks. Cantonal banks and regional banks, however, have distinctly lower claims and liabilities vis-à-vis non-residents and smaller business in foreign currencies than the corresponding groups of medium-sized or small banks. This mainly reflects the role of foreign banks in Switzerland not included in the traditional classification. Nonetheless, the three traditional groups of Swiss banking can be regarded as reasonable proxies for bank size. The balance sheet total of the average Big bank was roughly 15 times the corresponding figure of the average cantonal bank and more than 70 times the figure of the average regional bank in February 1987.

3. The empirical investigation

In this section, the response of Big banks, cantonal banks, and regional banks to monetary policy is examined. We proceed as follows. First, the data is described and presented in form of several charts. Then, the VARs are introduced and the impulse response functions are examined.

Table 1

Balance sheets of Big banks, cantonal banks and regional banks in February 1987

	Big Banks	Cantonal banks	Regional banks
Number of institutions	5	29	37
Average balance sheet total in million SFr	86,296	5,337	1,180
Group's share of total balance sheet total in %	68	25	7

	Share of group's balance sheet total in %				Share of group's balance sheet total in %		
Assets	**Big Banks**	**Cantonal banks**	**Regional banks**	**Liabilities**	**Big Banks**	**Cantonal banks**	**Regional banks**
Cash	6	3	3	Sight deposits	18	9	11
- of which Sfr. claims on residents	2	2	2	- of which Sfr. claims on residents	5	7	10
Interbank time deposits	27	14	6	Time deposits	44	15	7
- of which Sfr. claims on residents	2	11	6	- of which Sfr. claims on residents	13	15	7
Securities	12	7	9	Savings	14	37	41
- of which Sfr. claims on residents	4	7	8	- of which Sfr. claims on residents	12	36	40
Unsecured loans	13	5	6	Bonds	12	30	32
- of which Sfr. claims on residents	5	4	6	- of which Sfr. claims on residents	12	30	32
Secured loans	30	64	69	Capital plus reserves	6	4	5
- of which Sfr. claims on residents	25	63	69	- of which Sfr. claims on residents	6	4	5
Other assets	11	7	6	Other liabilities	6	4	3
- of which Sfr. claims on residents	4	6	6	- of which Sfr. claims on residents	4	4	3
Total	**100**	**100**	**100**	**Total**	**100**	**100**	**100**
- of which Sfr. claims on residents	41	93	97	- of which Sfr. claims on residents	51	96	97
- of which claims in other currencies on residents	8	2	1	- of which claims in other currencies on residents	9	1	1
- of which Sfr. claims on non-residents	16	4	1	- of which Sfr. claims on non-residents	7	2	1
- of which claims in other currencies on non-residents	36	0	0	- of which claims in other currencies on non-residents	32	0	0

Note: See annex for data description.

Data

The balance sheet items used in the empirical part of this study are *deposits, loans, cash and securities*. All series refer to Swiss franc claims and liabilities to residents. The series were constructed as described in the preceding section, and then deflated by the CPI to obtain real variables (in 1993 prices). Deposits consist of sight deposits, time deposits, and savings deposits. Loans can be divided into secured loans and unsecured loans. Cash includes currency and sight deposits with the SNB, the PTT, and other banks. Securities are bills and bonds.

Graph 1 shows the evolution of real deposits for Big banks, cantonal banks, and regional banks, both in levels and annual growth rates. All three bank groups reveal the contractions in monetary policy in 1981/82 and after 1989. Real loans are shown in Graph 2. Swings in bank lending are largest for the Big banks, followed by the regional banks and the cantonal banks. The turning points as measured by the peaks and troughs of annual growth rates indicate that loans follow deposits with a lag. Graph 3 shows the evolution of the three bank group's liquid assets. In comparison with deposits and loans, the liquid assets display larger swings. This goes for all bank groups and for Big banks in particular.

Some basic economic indicators of the Swiss economy for the period under review are summarised in Graph 4. There were two marked accelerations and decelerations in inflation (preceded by equally marked accelerations and decelerations in money growth; see Graph 1). Output growth dropped both in the early 1980s and in the early 1990s. Low output growth in 1978 can be attributed to the then exchange crisis. The exchange rate along with structural problems may also have played a role in delaying the recovery in the 1990s.

The indicator of monetary policy used in this study is a short-term interest rate (three-month Swiss franc deposits in the euromarket). This variable was chosen although it is not controlled by the SNB. The traditional monetary indicator of the SNB, the monetary base, is passed over because two innovations caused a significant shift in the demand for base money in 1988: the new system for reserve requirements and the gradual introduction of a new interbank clearing system (SIC). Thus the monetary base does not provide reliable information on the pre-history of the recession in the early 1990s.

Model

The empirical strategy is to estimate small unrestricted VARs for each bank group, and to examine the impulse response functions for monetary policy shocks. We consider five-variable VARs, consisting of the interest rate, real GDP, real deposits, real loans, and real cash and securities. We also present evidence for three-variable VARs described in the next section. Each VAR includes four seasonal dummies.[5] Data are quarterly as no monthly figures are available for output. All variables are in levels and (except the interest rate) in log-form. Lag lengths are determined by a sequence of encompassing tests performed in PcFiml 9.0. The estimation period is 1977:2-1996:3.

The unanticipated changes in monetary policy are identified by decomposing the residuals in a triangular fashion, the Choleski factorisation. Using this standard orthogonalisation method forces a causal structure on the system. Thus, the ordering of the variables may affect the results. We place the policy variable – the short-term interest rate – first. This implies that the policy variable has a contemporaneous effect on output and the banks' portfolio, but responds to output and the banks' portfolio only with a one-period lag.

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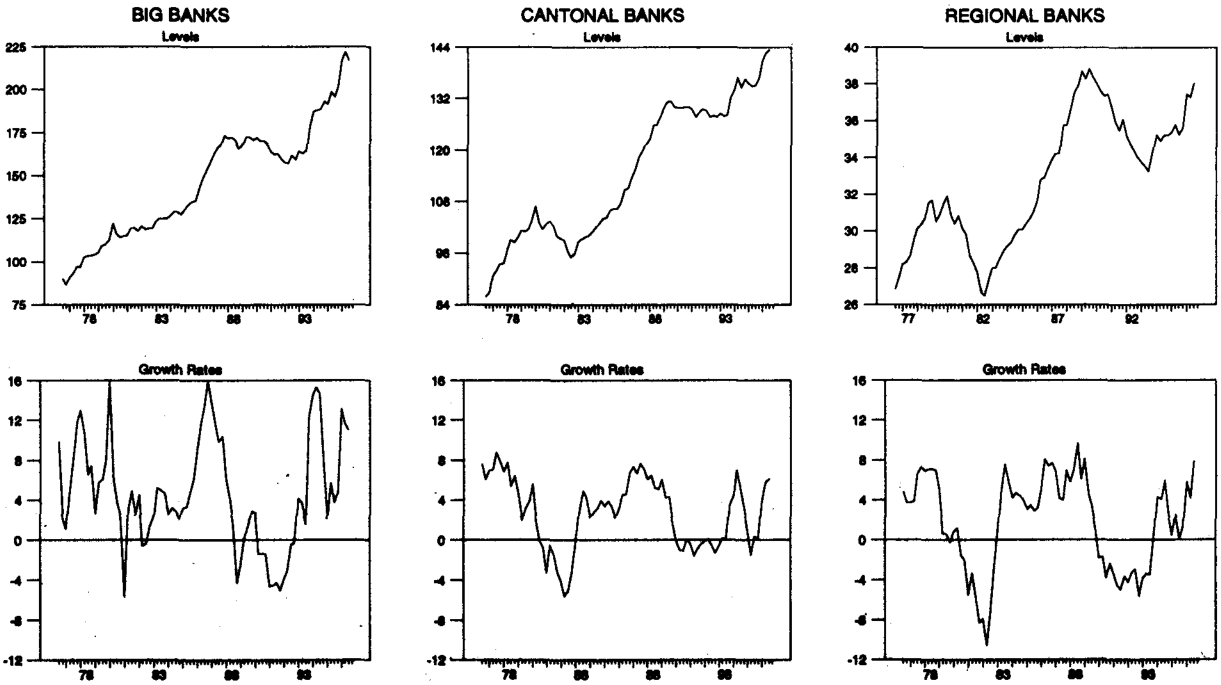
5 In addition, a dummy for the introduction of the SIC and the new reserve requirement system in 1988:1 is included in all VARs with cash and securities, and a dummy for the adoption of new accounting rules in 1995:1 is included in all VARs with loans.

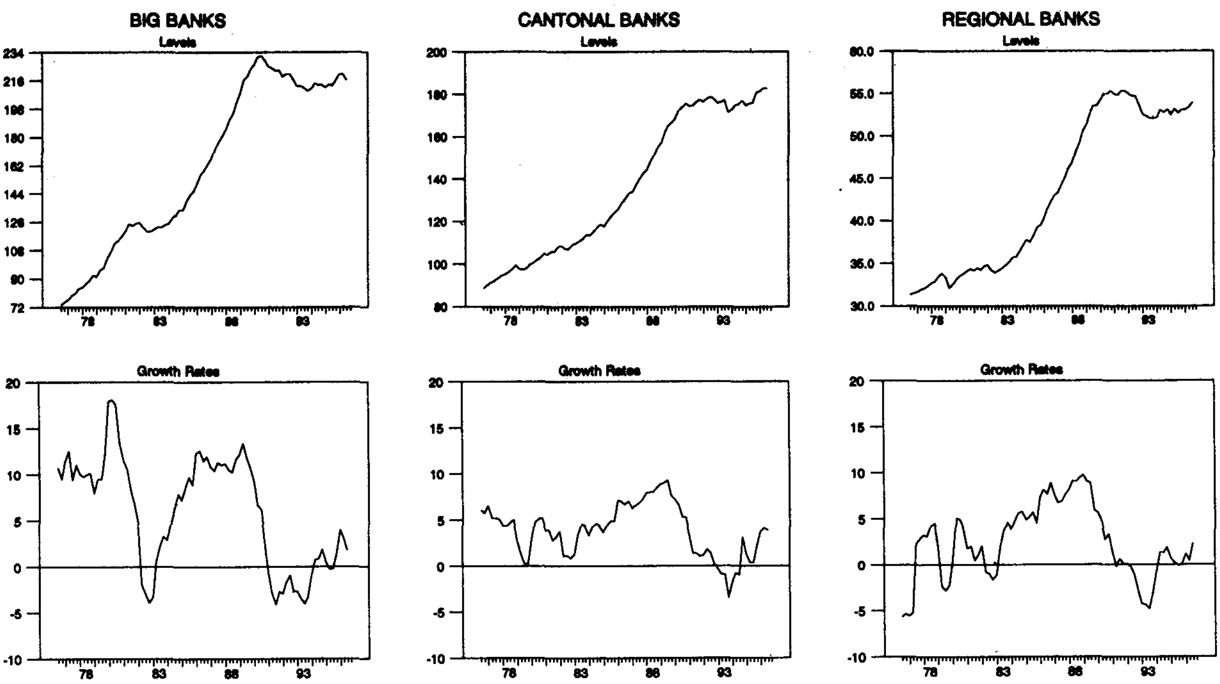
Graph 1

Real deposits



Graph 2

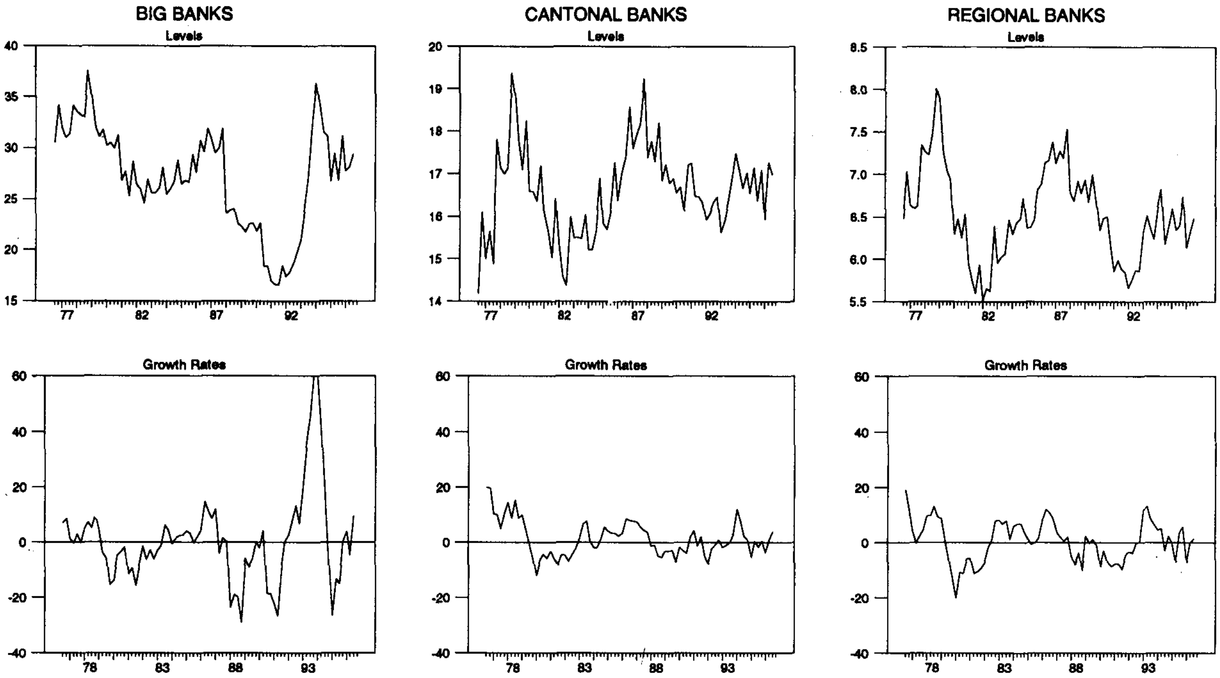
Real loans



Note: Quarterly data. Levels in billions of 1993 Swiss francs. Growth rates in year-on-year percentage changes.

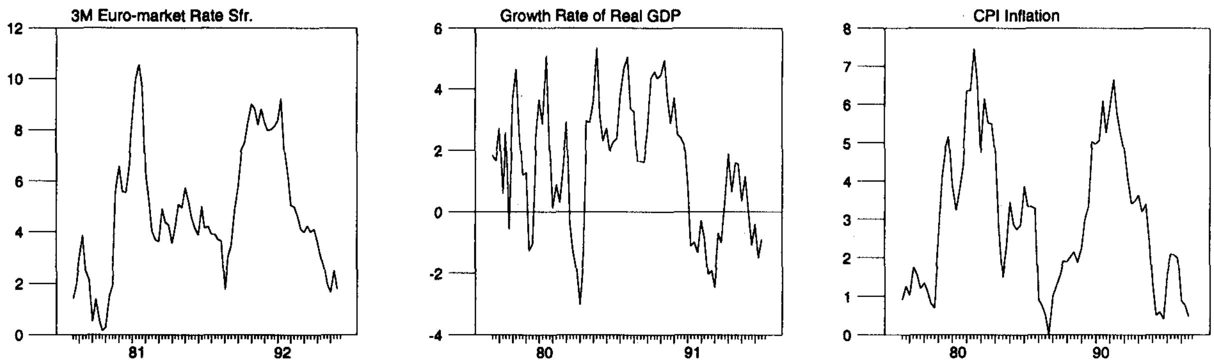
Graph 3

Real cash and securities



Graph 4

Interest rates, output growth and inflation



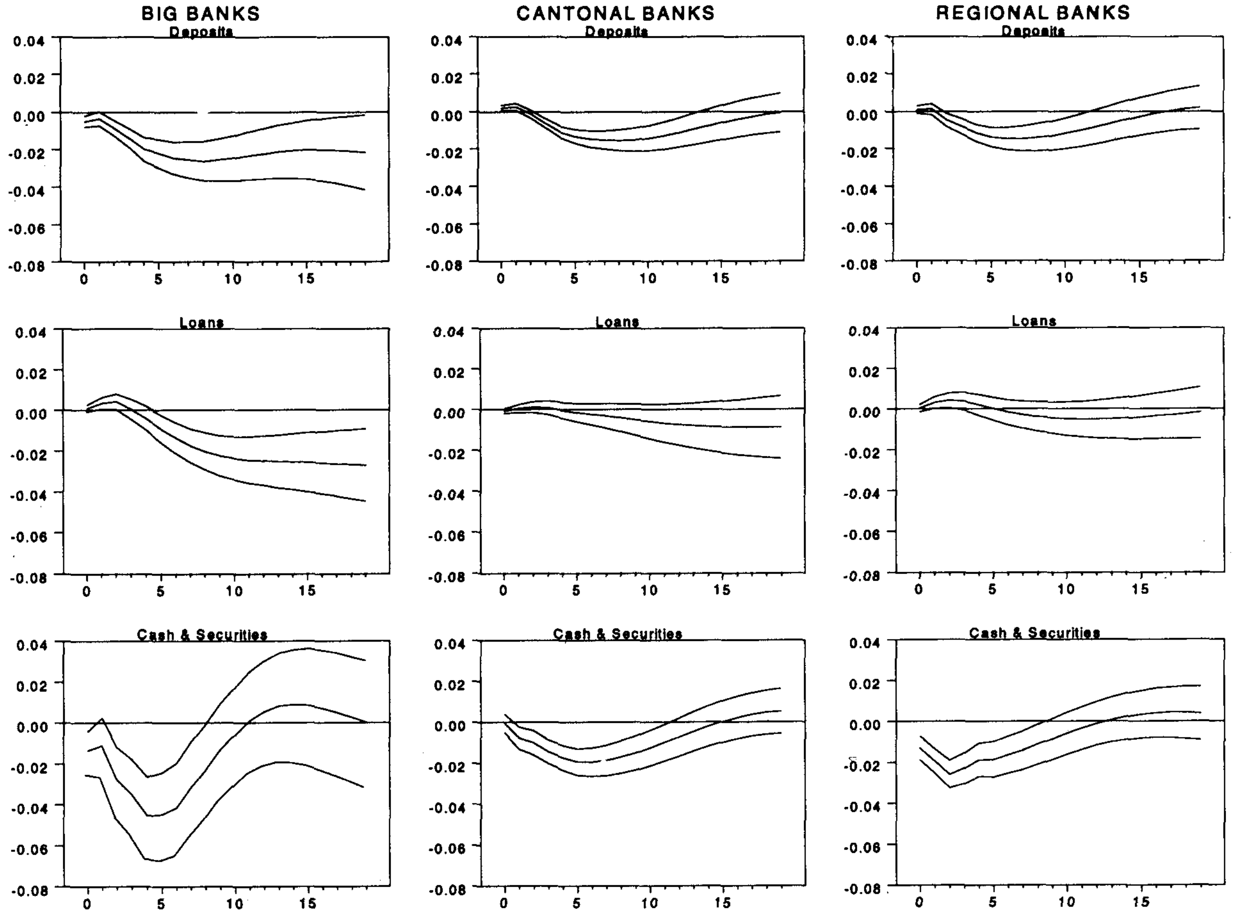
Note: Quarterly data. Interest rate in percent. Levels of other variables in billions of 1993 Swiss francs. Growth rates in year-on-year percentage changes.

Results

Graph 5 summarises the results for the five-variable VARs based on the impulse-response functions. In each panel the solid line plots the response to an interest rate shock which is normalised to correspond to a one percentage point increase in the current rate. The dashed lines indicate the plus and minus one standard deviation band of uncertainty associated with the estimate. The band is generated from Monte Carlo simulations with 1,000 draws. All responses are shown over a 20-quarters horizon.

Graph 5

VAR-5: responses to interest rate



Note: VARs with 5 variables (interest rate, output, deposits, loans, and cash and securities). Twenty-quarter response of variables to an interest rate shock (3-month euromarket rate Sfr.). Sample period is 1977:2–1996:3. Overall lag length is 3.

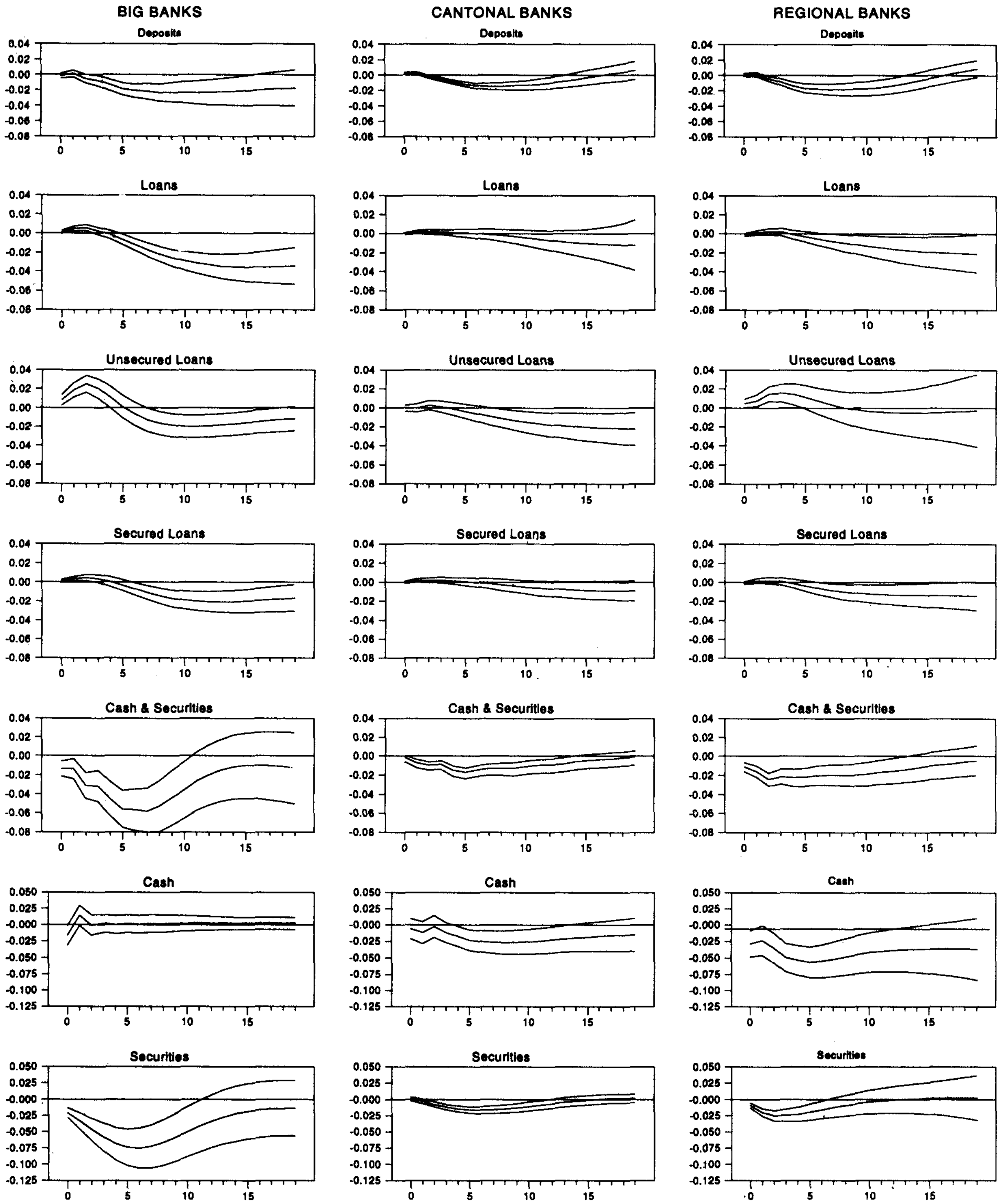
First, consider the effect on deposits portrayed in the first row of the graph. For all three bank groups we observe that the deposits decline after an interest-rate shock. The impact is similar for both cantonal banks and regional banks, while Big banks suffer somewhat larger losses in deposits. The evidence on the cross-sectional differences is weak however, because the one-standard-deviation bands around the point estimates overlap.

Turning to the impact of an interest-rate shock on loans, we observe that loans first rise and then decline. Notice, however, that the one-standard-deviation bands include the zero line for cantonal banks and for regional banks. A possible explanation of the 'perverse' short-term effect is that the stock-building of firms in the early stage of a recession may lead to a higher demand for bank loans. The subsequent decline of loans starts later than that for deposits and is more persistent. Based on point estimates, we find that Big banks display the largest response of loans among the three bank groups.

The effects on the banks' liquid assets are summarised in the third row of the graph. We find that an interest-rate shock triggers an immediate decline in the cash and securities at all bank groups. Again, the effect is largest for the Big banks. Differences between cantonal banks and regional banks once again are negligible.

Graph 6

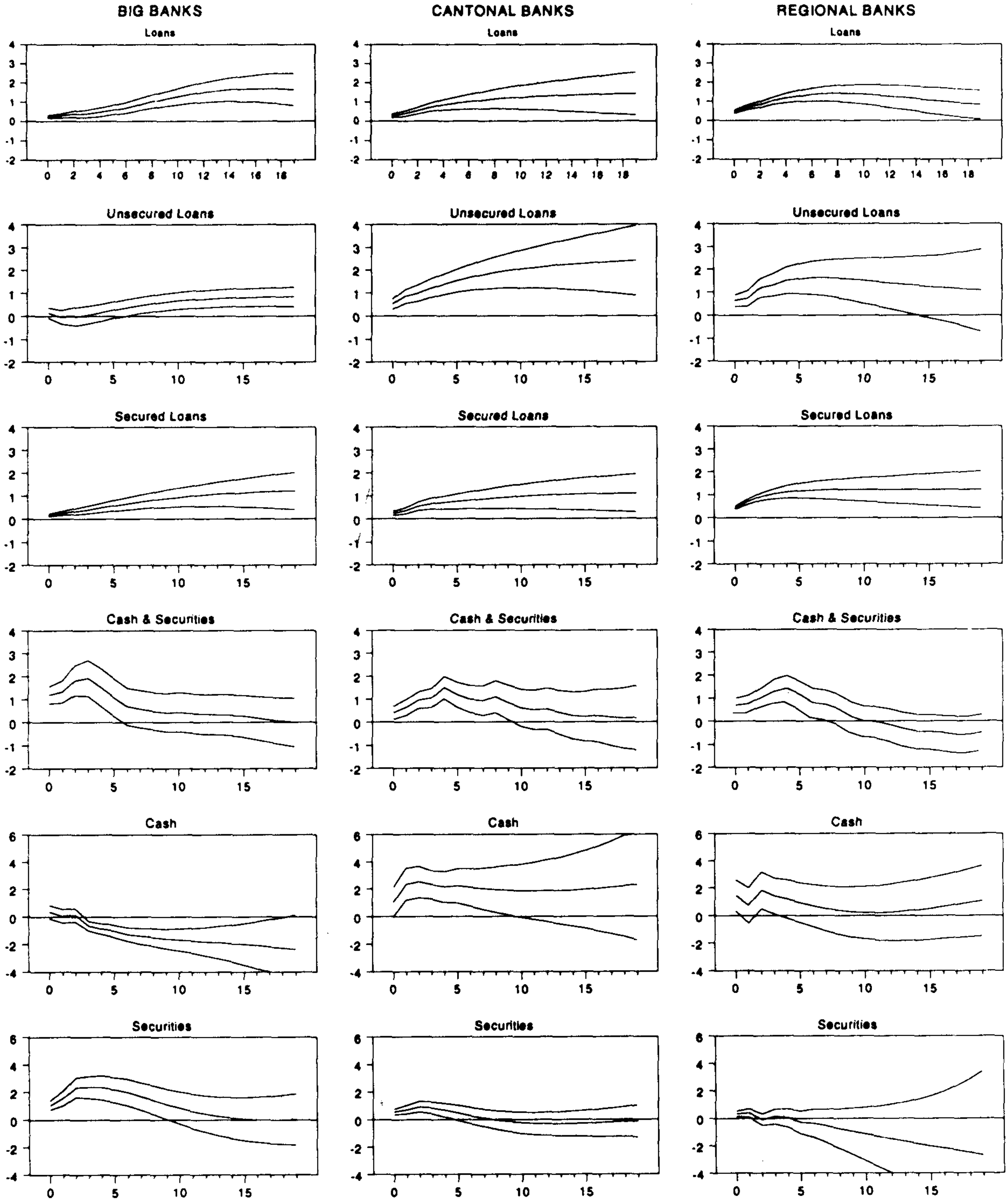
VAR-3: responses to interest rate shocks



Note: VARs with 3 variables (interest rate, output, and balance sheet position). Twenty-quarter response of variables to an interest rate shock (3-month euromarket rate Sfr.). Sample period is 1977:2–1996:3. Overall lag length is 5 for all VARs in rows 1, 2 and 5; 3 for VARs in rows 3, 4 and 7; and 2 for VARs in row 6.

Graph 7

VAR-3: responses to deposit shocks



Note: VARs with 3 variables (interest rate, output, and balance sheet position). Twenty-quarter response of variables to an interest rate shock (3-month euromarket rate Sfr.). Sample period is 1977:2–1996:3. Overall lag length is 5 for VARs in row 4; 4 for VARs in row 1; and 3 for VARs in rows 2, 3, 5 and 6.

The overall picture offered by Graph 5 is in line with the thrust of results from the literature on the bank lending channel. The decline in bank loans triggered by a monetary contraction lags the fall in banks' deposits and liquid assets. The cross-sectional implications of the bank lending channel, however, are barely borne out by the data. In particular, we do not find the response of loans to be largest for regional banks and smallest for Big banks as suggested by the predictions of Kashyap and Stein (1995).

A more detailed account of the impact of interest-rate shocks on bank balance sheets is provided by Graph 6. The graph summarises the response of individual components of loans (unsecured loans and secured loans) and liquid assets (cash and securities) to a monetary contraction. This extension is motivated by two observations. First, the composition of the loan portfolio varies across bank groups, possibly reflecting the relative importance of mortgage loans. Second, the demand for cash gradually shifted downward in 1988 and 1989, when the SIC was introduced and the SNB adopted a new system for reserve requirements. The impulse response functions shown in Graph 6 are based on three-variable VARs, consisting of the interest rate, output, and the balance sheet position in question. The number of variables in each VAR was reduced from five to three variables to have sufficient degrees of freedom. Notice, that the 21 panels of Graph 6 are based on the impulse response functions from 21 different VARs (while the nine panels of Graph 5 were from three different VARs).

The results for deposits, loans, and the total of cash and securities are in line with those from the five-variables VARs portrayed in Graph 5.[6] In addition, we find that the striking response of Big bank cash and securities is largely borne by the securities portfolio. We do not find much difference between the responses of unsecured loans and secured loans. Yet unsecured loans display a more pronounced temporary rise in the first few quarters after the interest rate shock.

We next focus on the impact of a shock in the bank group's own deposits on loans and on cash and securities. This is done because the response of deposits to an interest-rate shock differs across bank groups. Since we measured the effect of an interest-rate shock to be largest for Big banks, it is not entirely surprising that Big banks display the largest impact on loans or cash and securities too. We adopt the same strategy as for Graph 6, except that the bank group's own deposits take the place of the interest rate as the policy variable. The impulse response functions are computed from a normalised one-standard-deviation shock to deposits.

The responses displayed in Graph 7 show some interesting patterns. Overall, they are the most favourable in this paper in terms of sheer consistency with the predictions of the theory. In particular, unsecured loans of Big banks now respond relatively slowly to a monetary policy shock. At the same time, the pronounced and rapid response of Big banks' securities is still recognisable. Both patterns are consistent with the predictions of Kashyap and Stein (1995). Unfortunately, the results seem to be less robust to changes in the overall lag length and the ordering of the variables (the monetary policy variable placed last) than the results from the impulse response functions for interest rate shocks.

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6 The responses to interest rate shocks do not depend critically on the ordering of the variables. We also examined several alternative specifications of the VARs reported in this paper to check the robustness of the results. In turn, we replaced (i) variables in levels by first differences, (ii) real variables by nominal variables, and (iii) quarterly data by monthly data (where monthly output figures were approximated with a spline function). The main results are not affected by these changes.

Concluding remarks

In this paper, we have examined the response of bank portfolios to monetary policy shocks across various bank groups in Switzerland. The hypothesis, which stems from Kashyap and Stein (1995), predicts that a contraction of monetary policy causes a relatively strong reduction of small banks' loan portfolio and of large banks' securities holdings. Based on the point estimates of the impulse response functions, we find that the responses are consistent with the predictions for securities but not for loans. In contrast to the theory, Big banks seem to have the strongest decline in loans after an interest rate shock. The evidence is weak, however, if the uncertainty of the estimates is taken into account.

A number of possible problems should be noted. One is the short sample period. The sample period covers only two distinct episodes when the SNB deliberately tightened monetary policy to bring down inflation. This is probably not enough for an investigation of cross-sectional differences of bank group responses to monetary policy.

Another problem is the measure of monetary policy. We have identified the monetary innovations based on a standard Choleski factorisation in an unrestricted VAR. Various authors have imposed more structure on the VAR and have proposed alternative identification schemes.

The third problem is the role of the Big banks. Switzerland's Big banks hold a large share of their balance sheet total as claims and liabilities vis-à-vis non-residents or in foreign currencies. This makes the asset and liability management more complex, and we may have a problem of omitted variables.

The fourth problem results from the observation that Big banks probably have a different structure of customers than the cantonal banks or regional banks. It raises the possibility that differences in the response of bank portfolios to a monetary policy shock are driven by differences in demand. This is, of course, the identification problem mentioned earlier. The solution proposed by some authors is to dig down to the level of the individual loan contract. Such a data set is not available, however. In their 1997 paper, Kashyap and Stein find an intermediate solution and examine the bank lending channel based on a large panel data set consisting of the quarterly balance sheets of virtually all US banks. A comparable exercise for Switzerland could be made for the post-1987 period only. For a study with a longer time horizon, there is no escape from the semi-aggregated figures for Big banks, cantonal banks and regional banks used in this paper.

Annex: data sources

Balance sheet variables

Nominal monthly series for 1975:07 to 1987:02 from the SNB-EASY-Databank. Construction of the corresponding series for 1987:02 to 1996:11 as described in the text based on the balance sheets from individual banks (SNB-IPSO-Databank).

In the empirical section all balance sheet variables refer to claims and liabilities in Swiss francs vis-à-vis residents; all variables are quarterly and (except interest rates) are in logs and deflated by the CPI (1993=100). The composition of the balance-sheet variables is as follows (see *Bankenstatistisches Beiheft, SNB Monatsbericht*, A10, A11, A20, A21):

Cash	*Kasse, Giro- und Postcheckguthaben + Bankendebitoren auf Sicht.*
Deposits	*Bankenkreditoren auf Sicht + Bankenkreditoren auf Zeit + Kreditoren auf Sicht + Kreditoren auf Zeit + Spareinlagen + Depositen- und Einlagehefte.*
Loans	Unsecured loans + Secured loans.
Secured loans	*Kontokorrent-Debitoren mit Deckung + Feste Vorschüsse und Darlehen mit Deckung + Hypothekaranlagen.*
Securities	*Wechsel und Geldmarktpapiere + Wertschriften.*
Unsecured loans	*Kontokorrent-Debitoren ohne Deckung + Feste Vorschüsse und Darlehen ohne Deckung.*

Other variables

Interest rate	Interest rate on 3-month Swiss franc deposits on the euromarket
Output	Real Gross Domestic Product

All data are from SNB-EASY-Databank, except where indicated otherwise.

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Comments on "Monetary policy, aggregate demand, and the lending behaviour of bank groups in Switzerland" by Olivier Steudler and Mathias Zurlinden

by Koichiro Kamada*

Introduction

I have found that this is an impressive paper on lending channels of monetary policy transmission. As Kashyap and Stein (1994) state clearly, monetary channels and lending channels of monetary policy transmission are not exclusive. In the monetary channels view, there are two assets: money and bonds. Prices are assumed to be sticky and so an inflation rate is to be almost zero. When monetary authority reduces base money, money supply declines in nominal and also real terms. Nominal interest rates of bonds rise and so do real interest rates since the inflation rate is zero. High real interest rates mitigate private final demands and thus real economic activities. Advocates of lending channels (e.g. Bernanke and Blinder (1988)) claim existence of another source that affects real activities of the economy. They consider three assets: money, publicly issued bonds and bank intermediated loans. If bank loans are special in some sense, monetary authority can influence the real economy by controlling an amount of bank loans. A critical point is whether bonds and bank loans are imperfect substitutes. In order that lending channels of monetary policy transmission exist, both firms and banks take the two assets as imperfect substitutes.

Steudler and Zurlinden examine banks' imperfect substitutability between loans and bonds, using data of the Swiss banking sector. They use data disaggregated into three categories in size and explore the Kashyap and Stein hypothesis that small banks change loans more rapidly than large banks. They have found that the opposite is true in the Swiss banking sector.

My comments are constructed as follows. In Section 1, I first present theory that supports the Kashyap and Stein hypothesis. I then summarise the authors' empirical procedure and main result and present my interpretation of the main result. I claim that their result does not necessarily deny existence of lending channels. Finally, I also point out an interesting fact in the Swiss data. In Section 2, I list some identification problems that should be noted in finding lending channels. In Section 3, I present theory of firms' liability choice to complete lending channels. In the final section, I make some comments and questions on the econometric methods taken by the authors.

1. Reaction of banks against monetary policy

A bank's liability choice depends on its default risk (or a risk premium). Faced with a reduction in reserves, banks switch from deposits to liabilities that require less reserves (CD, CP, or equity). Kashyap and Stein (1994) argue that small banks may suffer from larger default risk than large banks. Hence, responding to reduction in reserves, small banks reduce loans rather than issue non-deposit liabilities. So monetary policy transmit more effectively through small banks. This can be called a default risk hypothesis of bank liability.

Using balance sheet data of the Swiss banking sector, Steudler and Zurlinden examine a variation of the default risk hypothesis by Kashyap and Stein (1995): small banks reduce their loans more rapidly than large banks during tight monetary policy. The authors have found that the

* The views in the comments above belong solely to Koichiro Kamada and not to the Bank of Japan.

hypothesis cannot be supported in the Swiss banking sector. That is, during tight monetary policy, large banks reduce loans more rapidly than small banks.

To begin with, the authors sort out the balance sheet data of the Swiss banking sector, based on size of banks. The authors claim that the traditional category of the Swiss banking sector (i.e. the Big, cantonal, regional banks) almost coincides with the grouping based on actual bank size after February 1987. This makes data available back to 1975. Next, the authors estimate unrestricted VAR with five variables: interest rates, real GDP, real loans, real deposits, and real cash and securities (level in log except for the interest rate). Using the estimated system, the authors examine impulse-response following an increase in the interest rate.

The authors have found that the Big banks reduce loans more rapidly than the cantonal and regional banks do, thereby denying the Kashyap and Stein prediction (1995). This is an impressive result. I am not sure, however, that the authors' result breaks the validity of the Kashyap and Stein prediction (1995) for the following reasons: (i) the authors use data concerned only with assets offered to residents. I am not sure whether this strategy is successful, since banks choose their portfolio, taking into account its total return and risk. So focusing on the partial data may generate misleading results; (ii) small banks do not reduce loans if the BIS risk-based capital requirements are already binding. So it is a natural result that loans by small banks react less to monetary shocks.

Moreover, the authors' result does not necessarily deny the existence of lending channels. Suppose that small banks deal with small firms. In this case, if small banks do not reduce loans, there are no effects through lending channels. If large banks also deal with small firms, however, there may exist lending channels.

Finally, using the Swiss data, I want to discuss the buffer hypothesis of liquidity asset holding that Kashyap and Stein (1994) find in the US banking sector. A bank's asset mix is a result of its optimal portfolio choice. To protect sudden withdrawal by depositors, banks have to keep some liquid assets. Liquid assets, however, earn low returns. Bonds are liquid but earn low returns; bank loans are illiquid but earn high returns. Kashyap and Stein (1994) argue that small banks have smaller shares of loans. The reason is that small banks need more liquid assets, since they are exposed to larger swings of deposits than large banks are. Table 1 shows us the validity of the buffer hypothesis by Kashyap and Stein (1994) in the Swiss data. On a residents-only basis, shares of liquidity assets (cash, interbank deposits, securities) are 8% for the Big banks, 20% for the cantonal banks, and 16% for the regional banks. So in a static sense, the Swiss banking sector has a similar tendency to that in the United States: the Big bank's share of liquidity assets is the smallest, the cantonal banks' is the middle, and the regional banks' is the largest. On an all-customers basis, similar shares are 45% for the Big banks, 24% for the cantonal banks, and 20% for the regional banks. This happens because assets against non-residents are held mostly as liquid assets. This may be specific to the Swiss banking sector. Graph 5 also shows us further evidence for the buffer hypothesis. Responding to a rise in the interest rate, deposits decrease first and then loans decrease. To balance banks' assets and liabilities, cash and securities decrease as a buffer.

2. Identification problems

As mentioned above, the authors do not necessarily deny lending channels. Then the next question is how to find lending channels. In this section, I discuss three points that should be noted in finding lending channels: (i) how to identify monetary policy stance, (ii) how to distinguish changes in loan supply from those in loan demand, and (iii) how to distinguish effects of lending channels from those of money channels.

First, I discuss identification of monetary policy stance. The authors use interest rates on 3-month Swiss franc deposits on the euromarket. A natural question is how well these interest rates reflect monetary policy stance of the Swiss National Bank. These rates, however, are likely to be exposed to shocks that the Swiss National Bank does not intend. Policy stance should be distinguished from other shocks on the interest rates. One way to do so is to specify an interest rate equation in the VAR model so that it reflects monetary policy stance. A useful specification is a variety of Taylor rules. This makes the models' interpretation easy. Note that somewhat ironically, if lending channels are so effective that monetary authority can control an amount of bank loans without raising any interest rates in the economy, monetary policy stance is hard to be identified.

Second, it is difficult to distinguish responses in loan supply from those in loan demand against interest rate shocks. Lending channels assume the former. A simple way of finding lending channels, though not satisfactory, is to see causality between loans and GDP. Real loans and GDP respond to interest rate shocks and react to each other in the following ways. GDP decreases due to a reduction in loan supply (the banking sector reduces loan supply) or loan demand decreases due to a reduction in GDP (the manufacturing sector reduces loan demand). Unfortunately, the authors present neither figures nor test statistics to show their relationships. So it is hard to see the causality. Another way of distinguishing changes in loan supply and demand is to see movements of close substitutes for bank loans. For instance, it may be inferred that loan supply is falling if an amount of CP is growing while that of bank loans is falling.

Finally, it is hard to calculate how much effects are attributable to money channels and to lending channels. As mentioned before, monetary channels and lending channels are not exclusive mutually. They work at the same time. When monetary authority tightens money, an amount of money shrinks, which in turn raises interest rates and discourages final demands. This is a monetary channel. Lending channels add further effects on final demands through banks' behaviour. So it is difficult to find how much is due to lending channels and money channels. This is true, even if the authors use disaggregated data of the banking sector. If the Big banks are dealing with more firms that are very sensitive to interest rate than the cantonal and regional banks are, loans of the Big banks respond more to interest rate shocks than those of the cantonal and regional banks.

3. Reaction of firms against reaction of banks

To complete lending channels, I have to show that changes in bank loans affect firms' economic activity. If firms can access alternative financial source, such as CP market, changes in bank loans have no effects on firms' economic behaviour. So existence of lending channels depends on firms' accessibility to alternative financial markets. Below I discuss why some firms stick to bank loans rather than issuing bonds publicly.

First, information of firms is often imperfect for ultimate lenders. Information of some firms is so imperfect that they cannot finance their projects without being monitored by banks. In this case, those firms rely on loans combined with banks' monitoring. Bonds cannot be substitutes for bank loans, since bonds are sold without monitoring requirements. Second, suppose that firms borrow from banks and are monitored. Then being monitored by banks carries information on return and risk of those firms. In this case, bonds cannot be substitutes for bank loans, since bonds lack such information. Finally, since monitoring is costly, firms cannot switch between loans and bonds costlessly. Firms stick to banks, once they get into long-term relationships with the banks. Note that the recent rapid growth of the non-bank financial sector will mitigate lending channels, although the share of non-bank intermediation is still small in most countries.

Intuitively, small firms suffer from asymmetric information so severely that they need to be monitored by banks and can hardly issue CP and corporate bonds. Thus, lending channels work more effectively through small firms. Assume additionally that small banks transact more frequently with small firms. Then if small banks reduce loans, lending channels work more effectively. Thus, correlation of size of firms and banks can play an important role in monetary policy transmission. Note that the effectiveness is reduced, however, if large firms issue CP and lend money to small firms.

4. Econometric methods

Finally, I want to make comments and questions on the econometric method taken by the authors. The first thing I note is on stationarity of variables used in the VAR models. The authors present no results of stationarity tests at all. So I hardly evaluate the legitimacy of the specification of their models. Since I have no data in hand so as to test their models myself, I use the information given in the paper and express my feeling. The authors specify VAR models, using level data of real GDP, real loans, real deposits, and real cash and securities. These variables usually grow at a positive rate, however. For instance, in Graph 1, the real deposits of the Big banks grow at the rate of about 7% annually, those of the cantonal banks at about 3%, and those of the regional banks at about 2%. The real loans have similar trends as seen in Graph 2. Stationarity of the real cash and securities is difficult to see in Graph 3. It is reasonable, however, to guess that they grow with other balance sheet items, such as the real loans and deposits. Graph 4 shows that GDP of Switzerland grows at the rate of 1-2% annually.

The impulse responses in Graph 5 are against the consensus among monetary economists that monetary policy can affect economic activity in the short term, but not in the long run. A plausible picture is that real deposits and loans increase on the Switzerland's balanced growth path and that against interest rates shocks, paths of deposits and loans deviate from the balanced growth path for a while, but revert to the path in some time. Examining Graph 5, I have found that an interest rate shock (a change in monetary policy) can have permanent effects on deposits and loans of the Big banks. Although similar effects on deposits and loans of the cantonal and regional banks may diminish away, the share of the cantonal and regional banks is only 32% (= 25% and 7%) in the combined balance sheet of all the banks in Switzerland. Non-stationarity of the models, if any, causes serious problems in interpretation of the models. So I recommend to reconstruct the VAR models by using detrended data of real loans and deposits

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