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THE ECONOMICS OF RECENT BOND YIELD VOLATILITY

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List of common symbols used in the tables

Δ	change (first difference)
|·|	absolute value
$\bar{R}^2$	adjusted R^2
SD	standard deviation
SE	standard error of the coefficient
SER	standard error of the regression
DW	Durbin Watson statistic
AR(x)	auto-regression of order x
Q(x)	p-value of Q-statistic; x is the number of lags
ARCH(x)	p-value of ARCH test; x is the number of lags
LM(x)	p-value of Lagrange-multiplier test; x is the number of lags
WHITE	p-value of White's test for heteroskedasticity
OLS	ordinary least squares
IV	instrumental variable
*	significant at the 10% level
**	significant at the 5% level
***	significant at the 1% level
(.)	figures in brackets under coefficient estimates are standard errors
–	not applicable
..	not available
R	(logarithm of the) bond yield
ΔR	change in the log of the bond yield times 100
IVB	implied bond yield volatility

AU	Australia
BE	Belgium
CA	Canada
DK	Denmark
FR	France
DE	Germany
IT	Italy
JP	Japan
NL	Netherlands
ES	Spain
SE	Sweden
UK	United Kingdom
US	United States

Shading in the tables indicates the preferred regressions or highlights specific results.

Introduction*

The extreme fluctuations in yields that swept bond markets in 1994 surprised traders, portfolio managers, treasurers and central bankers. Observers from countries that were enjoying low inflation for the first time in a generation had come to hope that their financial markets had seen the last of such bouts of volatility. Those from countries with strong records of price stability saw the volatility as puzzling or even as an undeserved and undesirable intrusion from abroad. Those who had welcomed rising bond prices and low volatility in 1992–93 as evidence of increased confidence in official commitments to price stability suddenly found themselves groping for different explanations for falling bond prices and disorderly markets.

These episodes of bond market turbulence have added force to the question, raised in some quarters, of whether markets have become too powerful. One way of posing this issue is to ask whether this volatility reflects economic fundamentals of inflation, growth and policy, or whether it represents a self-generated force sweeping across markets and bearing little relation to such fundamentals.

In fact, comparatively little is generally known about the forces driving volatility. Market participants may be years ahead of the journals, as Robert Merton has argued, but they rarely gather or disseminate what they know. To be sure, managers of options trading desks feed volatilities for various markets into the routines that evaluate their positions; and risk controllers track the value of these positions under a variety of scenarios. These activities lead to a great deal of analysis of the short-

* We would like to thank Martin Watts of J.P. Morgan for data and insight, Joseph Bisignano, Renato Filosa, Stefan Gerlach, Anthony Rodrigues, Greg Sutton, Frank Smets, Kostas Tsatsaronis and William White for their helpful comments, Henri Bernard, Angelika Donaubauer and Gert Schnabel for statistical assistance, Wilhelm Fritz for technical help, Stephan Arthur for preparing the graphs and overseeing the publication, and to Joyce Ogilvie for secretarial help. Any remaining errors are our sole responsibility.

term determinants of volatility. Nevertheless, much of what is learned remains proprietary. For their part, perhaps influenced by the demands of the financial marketplace, scholars have concentrated research on the very short-term estimation of the time series properties of volatility. Moreover, the October 1987 Crash's salience has channelled most of the work into the stock market, leaving some attention to the foreign exchange market. By contrast, the study of volatility in fixed income markets in general, and the bond market in particular, remains a substantial lacuna. Against this background, we aim to move the analysis of bond volatility out of the awkward corner in which it now stands.

This paper incorporates the insights of short-term analysis of volatility and then seeks to relate fluctuations in volatility over time and across markets to domestic economic factors and international influences. In particular, we set out to ground bond volatility in market participants' uncertainties regarding each country's inflation, growth, fiscal policy and money market yields. We also investigate the role of international factors, not only the spillover of volatility from one national market to another, but also the influence of increasingly mobile international capital flows.

We address this work in the first instance to central bankers, who have at least two stakes in the study of bond volatility. Central banks have set the task for themselves of extracting from such derivative instruments as options information regarding market participants' expectations and balancing of risks.[1] In addition, supervisors of financial firms at central banks and elsewhere need to understand the risks that banks are taking, including their positioning in volatility.[2] At the same time, we address market participants, on whose insights we have drawn but who may find useful our bringing together in one place analysis of different markets and times. And we hope that scholars of volatility recognise in this work their common understanding and find our study of the link between volatility and economic factors of some interest.

Our results can be summarised briefly. Building on the well-developed short-term analysis of volatility, we confirm that when bond volatility rises, it persists at high levels for weeks. We find that volatility generally responds more markedly to sell-offs in the underlying market than to rallies and suggest novel explanations for this regularity.

[1] See Bank for International Settlements (1994), Neuhaus (1995) and Bank for International Settlements (1996b), chapters V and VI.

[2] See Board of Banking Supervision Inquiry into the Circumstances of the Collapse of Barings (1995), pp. 67–77.

As regards domestic macroeconomic determinants, we argue that an economy's inflation record and expectations increase bond volatility over long periods and across countries. Fiscal imbalances, too, may have a similar effect. Near-term revisions in inflation or growth expectations, however, do not track the short-term evolution of volatility. At least in the extreme case represented by Italy, variations in market participants' apprehensions about the state of public sector finances appear to have some influence.

Volatile money markets in general make for more volatile bond markets. Such money market volatility can in principle be related to the *modus operandi* of monetary policy, but the precise nature of the link remains an important question for research.

On the international side, we find evidence of spillovers and of a powerful and hitherto unappreciated influence of foreign disinvestment. Spillovers gained in strength and geographical scope in the period of market turbulence in 1994. They tended to emanate from New York and, to a lesser extent, London. For Germany, France and Italy, we find that the rapid withdrawal of foreign investors was associated with a sharp increase in volatility.

The study is organised as follows. Section I explains the definition of volatility underlying the analysis and deals with issues of measurement. Much of the work is based on implied yield volatilities derived from options on government benchmark bonds for thirteen industrialised countries. Section II examines the time-series properties of bond yield volatility and its relationship with market movements. A key issue considered is whether volatility is directional, rising in periods of rising yields (falling prices). Section III then assesses the short-run and long-run link between bond yield volatility and fundamental domestic macroeconomic factors. The focus is on inflation, growth and fiscal policies. Section IV studies part of the link between bond volatility and monetary policy, through an evaluation of the relationship between bond yield and money market volatility. Section V then turns to the international dimension. It examines the international transmission of volatility and the specific role played by sales and purchases by non-residents. The conclusions summarise the main findings of the study and draw a number of policy implications.

Inevitably, a study of this kind relies on statistical and econometric evidence to substantiate some of the claims made. We have relegated the more technical aspects to extensive annexes.

I. Definition and measurement

1. Medium-term asset price swings and short-term volatility

Discussions of financial market volatility often confuse, and sometimes falsely separate, two related concepts: medium-term swings or misalignments in asset prices and short-term volatility, as traded in options markets. A common medium-term price swing extending over years is evident in the rise in bond market prices in 1992–93 and their decline through most of 1994 (Graph 1). By contrast, we use the term volatility to mean a summary measure of the size over some short-run of the percentage changes in bond yields, irrespective of their sign. More specifically, we adopt as basic measure the (annualised) standard deviation of daily percentage changes in yields.[3]

We recognise that for the most important functions of asset prices, such as setting a hurdle rate for an investment project, medium-term swings matter much more than short-term volatility. One can think in this connection, for instance, of the closing of manufacturing plants in the United States in the period of the strong dollar in the mid-1980s or of the investment boom in Japan on the back of the strong stock and real estate markets in the late 1980s.[4] The limited extant research on the effect of financial volatility on the macro-economy largely finds muted effects of short-term volatility but perhaps a stronger influence of medium-term swings.[5]

It is a mistake, however, to let the useful distinction between asset price swings and short-term volatility obscure the relationship between them. In particular, short-term volatility often sets in at the reversal of a fairly long swing in asset prices. High volatility marked the end of stock market booms in the 1920s and in 1987 in the United States and in the late 1980s in Japan. In bond markets, the generally high volatility that accompanied the price retreats in 1994 confirmed the link between the initial phase of bear markets and volatility spikes which had already been observed after the fall of the Berlin Wall in Germany and after price run-ups in 1986–87. In exchange markets, the dollar's drop in 1985–87

[3] See Annex I for the reasons why yield volatility was preferred to price volatility.

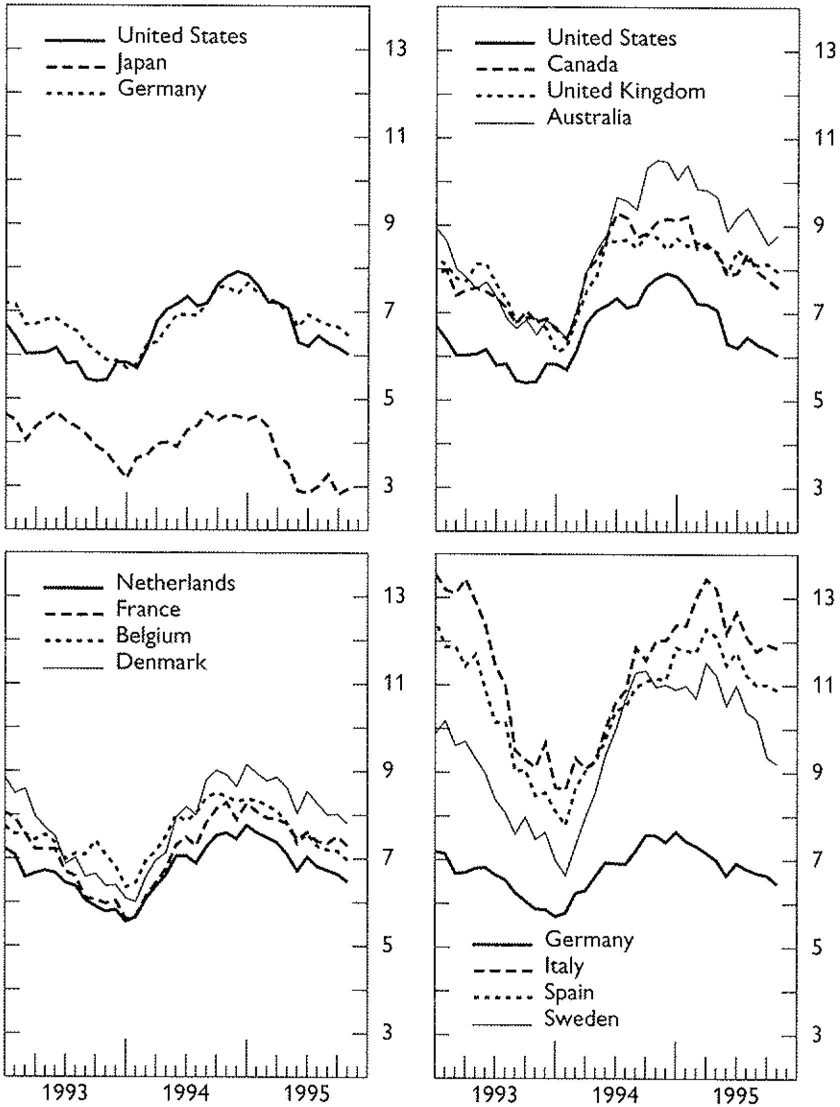
[4] See Inoue, Ishida and Shirakawa (1995) for the effect of perceived low capital costs in Japan, and McCauley and Zimmer (1994) for the effect on Japanese foreign direct investment.

[5] See International Monetary Fund (1984), Gagnon (1993), and Goldberg (1993).

Graph 1

Bond yields *

At end of month, in percentages



* In this and subsequent graphs and tables, yields on ten-year benchmark government bonds.

Source: Datastream.

following a five-year rise provided the occasion for repeated bouts of high volatility. One reason for this association may be that leveraged long positions accumulate in the bull period but are then run off rapidly as prices fall. Below we will provide some evidence on this point in connection with the 1992–94 bull-bear cycle in bonds.

2. *Implied versus historical volatility*

Settling on the basic unit of measurement, short-term yield volatility, still leaves open the issue of whether to construct the measure from the historical path of yields ("historical" or "realised"[6] volatility) or to obtain it from market prices of instruments that embody some view about volatility (the "implied" volatility of options on bonds). Our analysis relies primarily on implied volatilities.[7] However, owing to the short history of options markets and thus the limited availability of the data, it draws on the evidence regarding realised volatility to make observations about the long-term experience. In order to avoid confusion, it is important to understand the conceptual and empirical relationship between these two measures as well as the reasons for our choice.

Historical volatility is calculated as the standard deviation of *actual* percentage yield changes over a specific window, say, three-months. This is an ex post and rather mechanical measure. If one thinks of volatility as capturing the *uncertainty* about yield movements i.e. their random (unpredictable) changes, then it is not a priori clear how useful historical volatility can be.

Implied volatility, by contrast, is an ex ante measure. At one level, it can be thought of as a synthetic indicator of the market's view about the unpredictability of yield changes over the *remaining* life of the option's contract, in our case 3 months.[8] More generally, supply and demand in the market for options set the option's premium (price); and this price, together with prevailing interest rates, can be used to back out an implied volatility through an option pricing formula.[9] Supply and demand will

[6] In what follows, we use the terms interchangeably.

[7] More precisely, the basic database consists of weekly observations on implied yield volatilities for three-month over-the-counter (OTC) options on 10-year benchmark government bonds in thirteen major markets as quoted by a leading market-maker, J.P. Morgan. See Annex I for a detailed assessment of the merits of these data compared with implied volatilities derived from exchange-traded instruments.

clearly depend on a number of factors, notably the dealers' confidence in their own estimates and their willingness and ability to take positions. They will also be informed by the recent actual percentage changes in yields and corresponding *realised* volatility.[10]

While conceptually distinct, it is evident that the forward-looking implied bond yield volatility moves generally together with the backward-looking historical volatility as calculated over the immediate past (Graph 2). This is also true for stock and foreign exchange markets. The short-run deviations can be sizeable, but are bound to iron themselves out over time.

In sum, relying on implied volatility for the examination of the recent experience has the advantage of using a measure which is more closely tied to market uncertainty. At the same time, having to employ historical volatility for the longer-term analysis is unlikely to affect our basic conclusions given the observed relationship between historical and implied volatility.

[8] Another potential measure of volatility, straddling the two mentioned in the text, is that derived from econometric estimates which model the volatility process jointly with that of the variable whose movements are being predicted e.g. the percentage change in the yield. This is the family of Generalised Autoregressive Conditional Heteroskedasticity (GARCH) models of volatility. GARCH estimates resemble implied volatilities in that they are an attempt to derive an ex ante measure of uncertainty. They resemble historical volatility in that the information set on which the forecasts defining that uncertainty are made is generally very restrictive, most often consisting only of the past history of the variable itself. By contrast, implied volatility can be thought of as being based on all the information available to market participants. Moreover, in terms of the econometric analysis, implied volatilities are clearly superior to GARCH estimates since they allow a much richer set of questions to be addressed. GARCH techniques are data-hungry, having to rely on a very large number of observations in order to yield any sensible estimates at all (e.g. Figlewski (1994) and Engle and Mezrick (1995)).

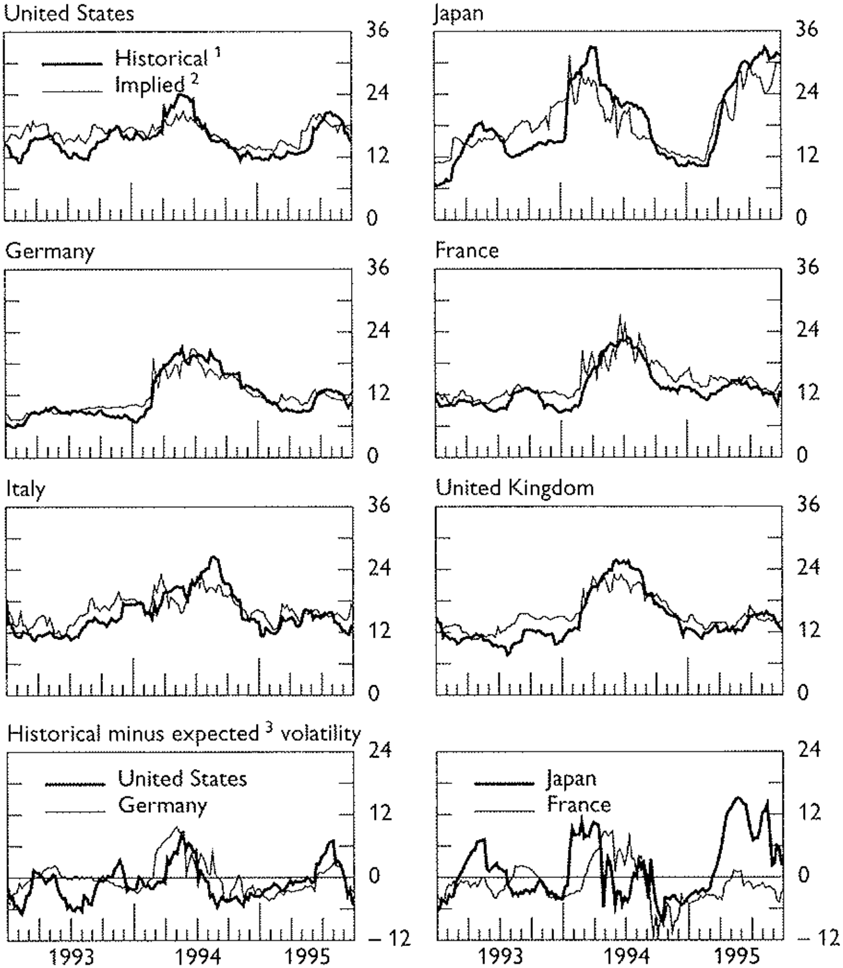
[9] Market-makers' methods for mapping premium prices into and out of implied volatilities vary somewhat across firms and across time, but the differences between pricing models are so subtle that market-makers find it convenient to quote their options in terms of the implied volatilities.

[10] It would clearly be wrong to argue that implied volatility is a mechanical reflection of realised volatility. Implied volatility is no more constructed from realised volatility than three-month LIBOR is constructed from recent overnight rates. A systematic investigation of the relationship between implied and realised bond volatility lies beyond the scope of the present paper. Such an analysis should also consider the predictive content of implied volatility for future realised volatility. Recent work on the US equity market suggests that both historical and implied volatility share some predictive power (Figlewski (1994)). Preliminary work on foreign exchange options suggests that implied volatility is a useful indicator (Bank of Japan (1995), Galati and Tsatsaronis (1996) and Jorion (1995)).

Graph 2

Bond yield volatility since 1993

In percentages



[1] Historical volatility is measured as the annualised standard deviation of daily percentage changes in bond yields calculated over the preceding ninety-one calendar days. [2] Yield volatility implied in three-month over-the-counter, at-the-money option contracts on ten-year benchmark government bonds, plotted at the time the contract is struck. [3] Expected volatility is implied volatility plotted at the time the contract expires so as to be aligned with historical volatility (e.g. the point in December is equal to the difference between historical volatility as plotted in December and implied volatility as plotted in September).

Sources: Datastream, J. P. Morgan and national data.

II. Properties of bond volatility: persistence, adaptivity and directionality

Research on volatility has generally responded to the needs of market participants for predictions of financial prices over very short horizons. This need has fostered more econometric development than insight into the economic forces underlying the dynamics of volatility. Moreover, little research has dealt specifically with the bond market. As a result, a review of what is known about bond volatility can be quite brief.

Prior econometric research suggests that, much as in the case of equity prices and exchange rates, bond yield volatility exhibits considerable inertia or *persistence*. In other words, when something pushes it up, volatility remains high for a considerable time and returns to its mean only gradually. Market participants argue that bond volatility is higher in bear than in bull markets, but to our knowledge their insight has so far not been taken up in the academic literature.

Our investigation of the time-series properties of volatility in the bond markets of industrial countries offers pervasive evidence of persistence. It also indicates that price jumps generally have an effect on implied volatility and that this effect is mostly asymmetric: often only sagging prices appear to have an impact. This asymmetry is analogous to that already amply documented for equity markets, but we suggest a number of potential economic explanations that differ from those put forward for that market. We find persistence harder to explain, at least in terms of the most widely shared academic paradigms.

1. *What goes up, comes down slowly*

The persistence of volatility must rank as its strongest, if not most profound, trait.[11] If this week's implied volatility is, say, twice its average, then next week's is quite likely to be about 90% higher than its average (Table 1). Thus, a surge in volatility decays at such a rate as to have a half-life of over a month. This trait accounts for a high fraction of the variance of volatility over time.[12] In particular, this week's volatility explains

[11] See Annex II for details, including an analysis of the issue of stationarity.

[12] The strength of persistence is the insight behind the univariate time series analysis of GARCH models, which asks a single time series to provide estimates of volatility and the relation of volatility to itself over time.

Table 1

Persistence of implied bond yield volatility[1]

	Persistence parameter[2]	$\bar{R}^2$	Sample begins on[3]
United States	0.90***	0.81	31.08.92
Japan	0.93***	0.87	31.08.92
Germany	0.96***	0.93	31.08.92
France	0.90***	0.81	31.08.92
United Kingdom	0.96***	0.92	31.08.92
Italy	0.84***	0.73	31.08.92
Canada	0.95***	0.90	31.08.92
Belgium	0.94***	0.90	31.08.92
Netherlands	0.97***	0.94	31.08.92
Spain	0.77***	0.58	16.11.92
Denmark	0.92***	0.83	14.02.94
Sweden	0.94***	0.89	14.02.94
Australia	0.88***	0.77	21.03.94

[1] Yield volatility implied in three-month over-the-counter at-the-money option contracts on ten-year benchmark government bonds. [2] Autoregressive parameter of AR(1) process estimated by OLS on weekly data. [3] The sample period ends on 22.05.95 for all countries.

Sources: J.P. Morgan and own elaboration.

anywhere from 58% to 93% of the variance of next week's volatility. Cross-sectionally, persistence appears to be noticeably shorter in the Spanish and, to a lesser extent, the Italian bond markets.

What does persistence signify? Those who have documented the persistence of volatility have taken little pain to make sense of it. At least three interpretations are possible: the pattern of news arrival, the digestion of news over time, and the memory of market participants.

The pattern of news

The first interpretation argues that volatility reacts immediately to the arrival of information ("news") but that the arrival of news itself exhibits persistence. News could follow hard on news, much as one hot day follows another in a heat wave.[13] But can news regularly arrive in a strong

[13] See Ito, Engle, and Lin (1990, 1992).

exponential declining pattern as required by this explanation? Perhaps the Kobe earthquake fits: the first flash of news, followed by smaller news shocks, followed by damage estimates of increasing accuracy. Yet it is at least as plausible that news and uncertainty build to crescendi as that they regularly diminish in decrescendi. A common class of counterexamples to the exponential declining pattern of information arrival is anticipated news events, such as national elections (see Box).[14]

Digestion of news

An alternative hypothesis is that information may arrive more or less uniformly in time but that market participants respond at various speeds, some immediately, others only with a lag, thereby generating persistence in volatility.[15] One can imagine a market as a pendulum at rest and news as a force that moves the locus of the tether. The weight on the pendulum would show something like an exponentially declining movement as it came to rest in its new position.

This explanation, however, is empirically unconvincing. Those who have watched the screen of US Treasury bond prices at the announcement of the monthly employment report at 8:30 Eastern Time on the first Friday of the month will find it hard to believe that diffusion or digestion takes weeks. There seems to be an exponentially declining pattern to volatility in the wake of news, but it takes seconds, minutes or hours, not weeks.[16] The prices of one-day at-the-money calls on the US 30-year bond, for instance, spike on the day before the first Friday of the month,

[14] Interestingly, social psychologists have argued for the reverse causation: in their view it is volatility that generates news, which in turn leads to persistence. In experiments, associating "explanatory" news (causal explanations or stories) with price movements leads to predictions that show greater persistence (slower regression to the mean) than those based on price movements presented with no story. The experimentor reads the evidence as suggesting that the media's understandable need to produce reasons for volatile market movements leads to persistence (Andreassen (1987)).

[15] See Kyle (1985) and Admati and Pfeiderer (1988), who argue that private information is only gradually incorporated into prices.

[16] "Real market prices of options are a forecast of bond market movement. These forecasts can be ...event-specific (newly released unemployment statistics were much higher than economists predicted). When such economic news is announced, all hell can break loose. The market is likely to have a *quick break* (fast change in prices) on the first words of the announcement and then correct just as quickly as additional information, such as a revision to the previous month's numbers, comes across the news tape. Volatility is worth paying up for on days when such action is possible." See Ray (1993), p. 198.

Box: **The volatility curve and the 1995 French elections**

The behaviour of volatility implied in bond option prices in the lead-up to the 1995 French elections illustrates the effect of anticipated new events. Such events provide counter-examples to the Kobe-style shock followed by ever-smaller after-shocks.

To see how options market participants priced in the approaching uncertainty, consider six options on each of six days in winter, spring and summer 1995. On each day, we have 2-week, 1-month, 2-month, 3-month, 6-month, and 1-year contracts. When the implied volatilities are plotted at the day of expiration of each contract, we trace a *volatility curve*. The solid heavy line, for instance, plots implied volatility for contracts traded on February 6th but maturing in 2 weeks, 1 month, and so on. This curve shows volatility at fairly normal levels over the immediate horizon but rising thereafter (Graph 3). Looking across the volatility curves, we see that volatility formed a hump that reached its highest in the contracts that matured after the election.

6th February: the 6-month contract showed the highest volatility. A 3-month put, for instance, would not provide insurance through the second round of the election, scheduled for 7th May.

6th March: the 3-month contract, which by then extended comfortably through the election, showed peak volatility.

20th April: the 1-month contract showed peak volatility on the eve of the first round of voting.

27th April: The one-month contract continued to show peak volatility, suggesting that uncertainty extended beyond the election (bracketed by the 2-week contract) to the choice of Cabinet by the new President.

4th May: the 2-week contract showed the highest volatility, just before the second round.

24th August: Political uncertainty continued to elevate volatility in near contracts, but six and twelve-month contract volatility had returned to the market norm of 12 percent.

Not all news arrive as a surprise and lingers.

but persist at high levels for no more than days.[17] Persistence over weeks seems too glacial to arise from market participants' response to news.

Memory

To at least one market participant, persistence reflects the memory of risk-averse traders:

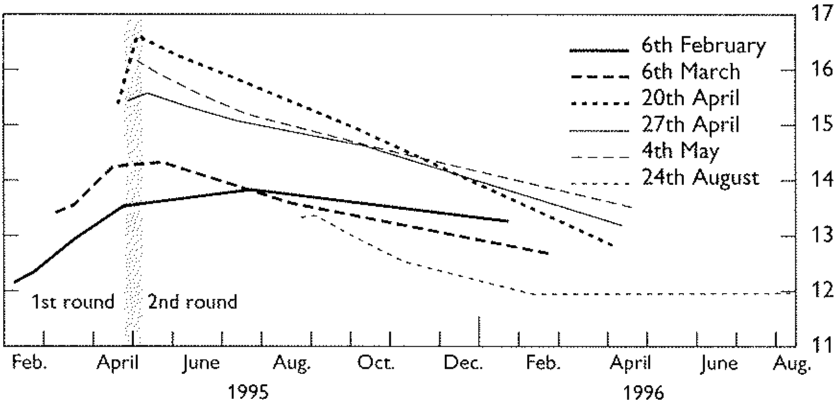
"After bond volatility on the Board reached 35% on October 19, 1987 (the long bond rallied *10 points* intraday), not many options traders were willing or allowed to sell short volatility. But look at December 1987. Although *historical* volatility was less than 10%, *implied* volatility

[17] See Ray (1993), p. 197. This observation regarding the biggest news of every month for the US bond market parallels observations of the stock market's digestion of the biggest regular company-specific news for shares. A study of the response of stock prices to specific earnings announcements found that the variance of price changes remains unusually high only for a matter of hours (Patell and Wolfson (1984)).

Graph 3

Volatility curves implied by French bond options on selected days in 1995

In percentages



Notes: Each curve plots at-the-money implied volatilities on the French ten-year benchmark government bond (OAT), using options with maturities of 2 weeks, 1, 2, 3 and 6 months and 1 year; option volatilities are plotted at their maturities. French voters cast their ballots in the first round on 23rd April and in the second round on 7th May, as indicated by the shaded area.

Sources: Hong Kong Shanghai Bank, Datastream and own calculations.

remained more than 15%. The memory of October was too fresh to allow much options selling."[18]

Indeed, implied volatility remained well above realised volatility for months after the Crash.[19] A similar pattern, with implied volatility overstating actual volatility across markets, can be observed in the bond markets in the latter months of 1994 (Graph 2). That buyers and sellers of options failed to anticipate the rise in volatility early in the year is not too surprising (bottom panel of the same graph). But that they then missed its reversion to the mean in the latter half of the year is more remarkable, given that gradual mean reversion is the best-known feature of volatility. Some form of risk aversion could be at work.

[18] See Ray (1993), p. 196.

[19] It has been noted that stock market volatility fell back relatively quickly after the 1987 crash but remained very high after the 1929 crash. Could the high implied volatility in November and December, 1987 reflect market participants' fear of policy errors that were not in the event made?

This hypothesis may go some way towards explaining the higher persistence of implied volatility compared with realised volatility following major painful shocks. It seems of limited value, however, in explaining the average persistence in realised volatility itself, also recognised to be quite high.[20]

We are thus left with somewhat of a puzzle: the most widely documented trait of volatility, persistence, is also one of the least understood.

2. *Cash market movements and volatility*

To say that volatility persists after it rises leaves wide open the question of why it rises in the first place. Analysts of volatility have argued that jumps in the cash market are likely to presage more jumpiness in the future. Here we investigate more broadly the empirical relationship between volatility and very short-term and medium-term market movements, focusing on the issue of asymmetries between upward and downward markets. We then turn to possible explanations of the phenomenon. While the evidence does not speak with one voice, it is broadly in favour of an asymmetric response of volatility to market direction, with price declines being associated with higher volatility.

Evidence of asymmetry

On the basis of the available long-term record of daily movements of the yield on benchmark bonds, bear markets look generally more volatile than bull markets (Graph 4). The dark-shaded periods of falling prices tend to show higher volatility than the adjacent lighter-shaded spells of rising prices. Volatility looks especially high at the beginning of bear markets.[21] The exception represented by the first Volcker bull market in US bonds in 1981–83 shows the power of the high volatility of money market rates at the time (see below).

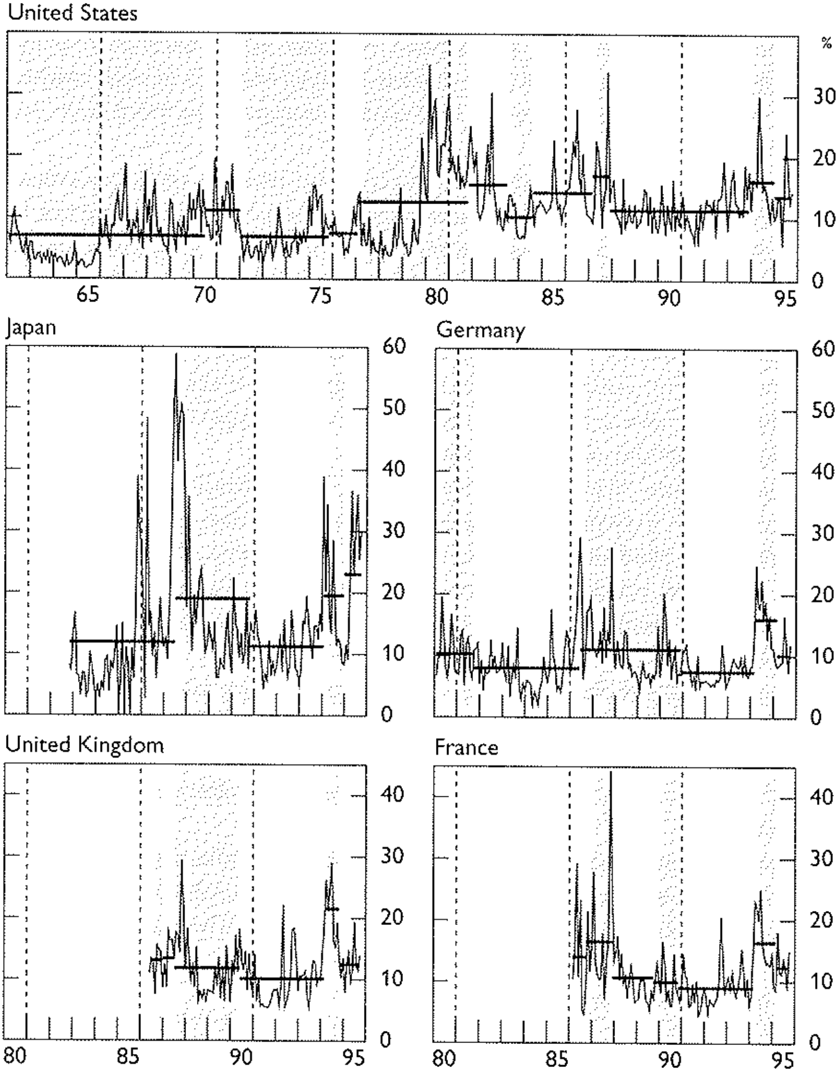
At the weekly frequency, our econometric analysis seeks to distinguish among at least four different cases in the relationship between implied volatility and market movements (Graph 5): (a) no relationship; (b)

[20] On this, see also Annex IV.

[21] Loeys (1994), p. 11, reports: "Looking over 10 of the major bond markets across the world over the last 9 years, bond yield volatility is on average ⅓ higher during bear markets than in bull markets. Out of 38 transitions between bear and bull markets, volatility was higher in the bear market during 32 transitions."

Graph 4

Bond yield volatility: a longer-term perspective



Note: Volatility is measured as the annualised standard deviation of daily percentage changes during calendar months in the yield on ten-year benchmark government bonds. The shaded (unshaded) areas represent bear (bull) markets and the horizontal lines the average volatility during these periods.

Sources: Datastream and national data.

symmetric (adaptive) relationship – yield changes in either direction raise volatility by the same extent; (c) semi-directional (asymmetric) relationship – only increases in market yields raise volatility; and (d) directional relationship – increases in yields raise volatility while declines depress volatility by the same extent. We regress implied volatility in each market on its level in the previous week and on the change in the yield over the previous week.[22]

Our evidence indicates that the dozen bond markets under examination fall into three categories (same graph). The largest bond market in the world, the United States, and its northern neighbour, Canada, show no statistically significant relationship between the implied volatility at the close of business on Thursday and the movement of bond yields from the close of the previous Thursday (Graph 5). This finding is a bit at odds with the longer-term visual evidence from Graph 4, where it seems that volatility *does* have a directional element. It may simply be the result of specific traits of the recent experience.

The Japanese market and, less clearly, the Spanish[23] and Swedish markets show a symmetric relationship between the direction of the market and implied volatility. This V-shaped relationship of movements of bond yields and volatility means that it is the *absolute* size of the change in the yield that feeds into the market participants' revision of volatility (Graph 5).[24]

The rest of the bond markets in the sample point to an asymmetric relationship between market direction and volatility, with market sell-offs associated with higher volatility and market rallies leaving volatility unaffected (Graph 5). In this respect, most bond markets resemble stock markets.[25]

Volatility responds fairly strongly to a sell-off in the underlying cash market. The estimated parameter varies between one-third and one-half in both the symmetric and the semi-directional cases. The outliers are the

[22]See Annex III for details.

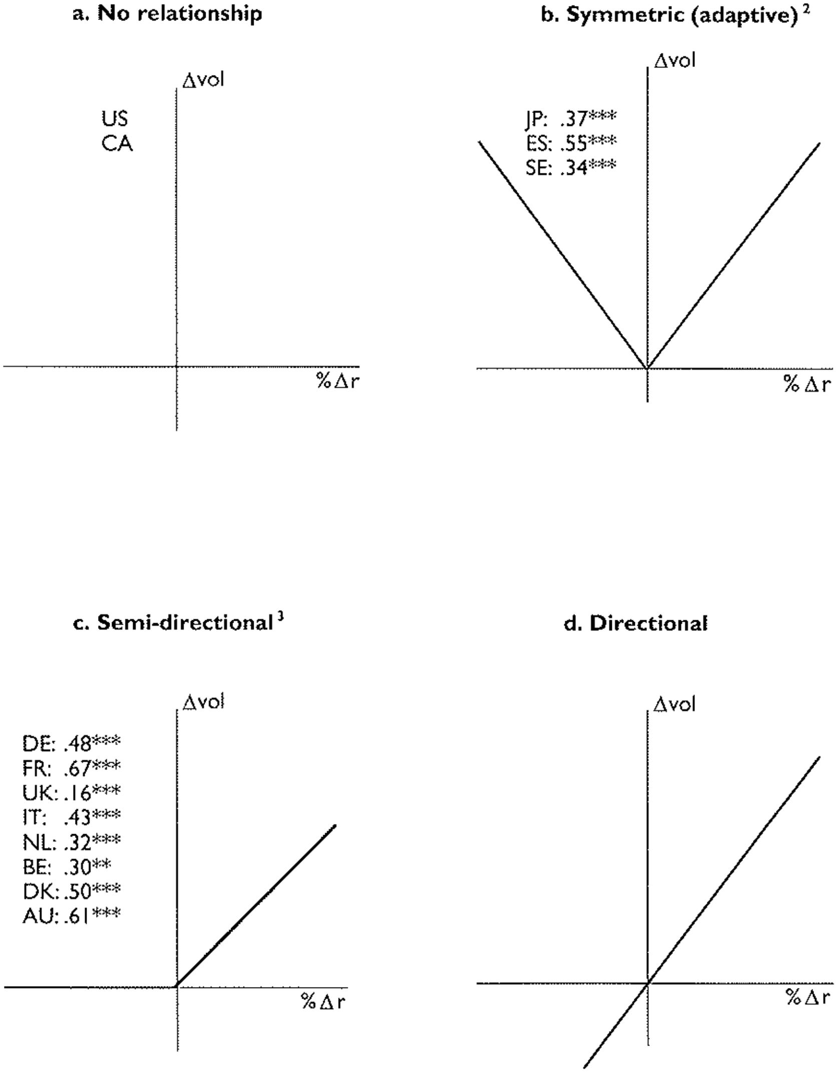
[23]But see Ayuso et. al. (1995) for evidence that conditional volatility based on realised rates shows an asymmetric relationship with market movements in the Spanish bond market.

[24]Karpoff (1987) argues for a somewhat similarly asymmetric relationship between the magnitude of a price change and the volume of trading. If the magnitude of the price change is associated with volatility, then putting the volume-price change relationship together with the price change-volatility relationship would be consistent with a volume-volatility link.

[25]See Rodriguez (1996) who finds a directionality for Germany, the United Kingdom and the Netherlands and no relation for Canada. However, he also finds no relation for Japan and France, and a directional relation for the United States.

Graph 5

Stylised relationship between implied bond yield volatility and changes in bond yields[1]



[1] Coefficient estimates of the suitably transformed weekly percentage change in the bond yield (first difference in the logs; Friday to Thursday) in an AR(l) regression for implied bond yield volatility. [2] Coefficients on the absolute value of the change. [3] Coefficients on positive changes only.

French, Australian and Spanish markets, where the response is particularly large. At the other end of the spectrum, the gilt market in the United Kingdom reacts comparatively little. Intriguingly, this evidence places it between the Continent and North America.

To get a sense of the size of the effect, consider a sell-off during one week that carries yields from 6% to 6.6%, or a change equivalent to 10% of the base yield. Taking the central range of one-third to one-half, such a sell-off would raise volatility by something between 3 and 5 percentage points.

Are the asymmetries related to medium-term market movements (bear vs. bull) or to very short-term, week-to-week jumps? We attempted to distinguish between these two possibilities using the weekly data on implied volatility since 1992. The period extends over one bull-bear-bull market sequence (Annex III). The results tended to confirm that the difference between volatilities in the two types of market is generally statistically significant, but they also suggested a stronger link between volatility and week-to-week movements.

Interpretations

A number of possible, partly complementary, hypotheses can be put forward to explain the link between market movements and volatility. These include traders' and investors' reactions to news, systematic variations in market-makers' willingness and ability to provide liquidity, leverage, specific trading strategies, and issuance patterns. We argue that leverage played a particularly significant role in the 1994 bond market turbulence.

Adaptive expectations – news impact curves? Academics have tended to interpret the link between market movements and volatility as reflecting reactions to "news". In the typical framework used to model volatility, the size of the forecast errors of the underlying variable (e.g. the asset return) is taken as the measure of news.[26] Accordingly, the term "news impact curve" has been coined to refer to high-frequency relationships similar to those portrayed in Graph 5 above.[27]

[26]This is the case of the popular and large family of GARCH models of volatility e.g. Bollerslev et. al. (1992) and Engle and Mezrich (1995).

[27]See Pagan and Schwert (1990), Engle and Ng (1993) and Hentschel (1995).

While conceptually consistent, this approach does not advance our understanding very far. Short-run asset returns or, in our case, percentage yield changes, are largely unpredictable given past history; there is thus a slight distinction indeed between market jumps and forecasts errors. Consequently, "news" is essentially just another name for the market movement itself. "Good" news is a rise in the market and "bad" news a decline. It is doubtful whether empirically all market movements represent reactions to news.[28] Moreover, in order to account for the observed asymmetries, an independent view about cognitive biases is necessary. Such biases may well be a trait of human behaviour, but we do not know whether other social disciplines have documented them.

Asymmetry of inflation risk. If a rise in bond yields is taken generally to represent the anticipation of higher inflation, then it is possible to put some more meat on the bare-bones explanation in terms of news. A macroeconomic interpretation can be constructed following Milton Friedman, who argued:

> A burst of inflation produces strong pressure to counter it. Policy goes from one direction to the other, encouraging wide variation in the actual and anticipated rate of inflation.[29]

As stated, this view implies an asymmetry, with higher inflation expectations (read yields) producing higher inflation (yield) volatility, but lower inflation (yields) not doing so. The view might be extended to suggest that where policymakers have a record of a measured and consistent response to higher inflation expectations, volatility might not rise in response to higher rates. Where policy-makers are likely to respond equally strongly to lower and higher inflation, the relationship might be symmetric.

This hypothesis may be appealing as a *general* explanation of the link between inflation, the level of yields and volatility (see below). It may also help to understand the historical association of high volatility with bear markets. And if sufficiently adapted it can also go some way towards explaining last year's events for several countries. For example, the

[28]For some evidence in the case of the stock market, see e.g. Cutler et. al. (1989). More generally, academics have coined the term "noise trading" to denote trading that is based on pseudo-information. See e.g. Black (1986) and Shleifer and Summers (1990).

[29]See Friedman (1976), p. 466.

symmetric relationship observed in Japan might be taken to suggest concerns about possible deflation too; the succession of expansionary fiscal packages and official rate cuts make this account plausible. Nevertheless, it still leaves certain cross-country puzzles unanswered. Why should no effect be observed in the United States and Canada and a strong asymmetry in Germany, given the latter's well-known successful record of low inflation? In addition, the hypothesis verges on hyperkinesis when applied to the relationship observed with *week-to-week* changes in yields.

Asymmetric responses by market-makers and investors. A second set of explanations relates to the wealth effects that result from changes in market prices and their induced impact on behaviour. Especially relevant in this context are systematic variations in *market makers' risk tolerance and capital* committed to making markets. Bear markets tend to impoverish market makers. At the level of the firm, poorer market-makers enjoy a thinner cushion between normal operations and the danger of bankruptcy and can therefore accept fewer risks.[30] Poorer market makers are also likely internally to shift more capital away from market making in bonds and into other activities, especially to desks trading instruments that have enjoyed a run of capital gains. In other words, downward markets can reduce the ability and willingness of market makers to provide liquidity. Their response during upward markets, by contrast, is likely to be less pressing and more muted. Note that these arguments apply better to the fairly low frequency relationship between bear markets and volatility than to the week-to-week relationship: managerial attitudes toward risk evolve and management committees allocate capital internally over months and quarters, rather than on a weekly basis.[31]

Leverage can greatly magnify the loss of liquidity due to market makers' reactions and can force ultimate investors to sell into falling markets. By leverage here we do not just mean debt-financing but also a mismatch between long-term (long duration) assets and short-term liabilities.

[30] See Greenwald and Stiglitz (1993).

[31] The response of market-makers to a price setback could accelerate at certain times in the compensation cycle. Traders just paid their bonuses might be reined in faster by managers who fear moral hazard: traders with little of their own money immediately at stake might take larger risks.

Higher yields reduce the value of the bond holdings while leaving the value of the floating-rate liabilities intact. As a result, the equity of leveraged investors falls and eventually they are forced to sell, precisely at a time when market makers are under pressure.[32] Moreover, if, as convincingly argued by Kindleberger and Minsky,[33] the *accumulation* of leverage can fuel asset price inflation, leverage also provides a link between medium-term swings in bond prices and the short-term volatility experienced in a bear market: the larger the price rise, the larger is the subsequent fall and hence the impact on liquidity and volatility.[34]

While much was said in the spring of 1994 about the influence of leveraged investors in bond markets, observers have offered few measures of the extent of leveraging and deleveraging. A partial attempt to fill this gap is reported in Table 2. The table assembles three kinds of evidence that point in the same direction. First, the body of the table shows the purchases of bonds by banks and securities firms. These purchases are taken as a sign of leverage because of the weight of short-term debt in the funding of these financial firms. The first memorandum item shows the interbank flows out of European money markets. One ultimate use of such funds would be to finance portfolios of European bonds held outside of the home markets.[35] The final two lines report the foreign holdings of Swedish and Spanish bonds financed with domestic repurchase agreements, that is, overnight or short-term funds. The evidence in the body of the table suggests that banks' and securities firms' leveraged positions were building up at a rate of something like $50 billion per quarter in the course of 1993. Then the process went into

[32]Our leverage effect differs from the popular explanation proposed by Fischer Black (1976) a generation ago for his observation that downward stock markets tended to be more volatile. He argued that as share prices decline, the (marked-to-market) debt-equity ratio of the corporate sector rises. This increase in leverage changes the character of equity, enhancing its option nature. Black's interpretation of the directionality of equity market volatility, of course, makes no sense when applied to bond markets. Our leverage effect refers not to a change in the character of the underlying security but to the limited tolerance for losses on the part of highly-geared investors.

[33]See Kindleberger (1978) and (1995) and Minsky (1982). For some evidence, see Borio et al (1993).

[34]The use of mark-to-market accounting combined with stop-loss strategies can exacerbate this effect. Re-setting the profit and loss balance to zero following a rise in prices reduces the cushion available to absorb declines in asset values without showing losses. This was reportedly a factor at work behind the turbulence in early 1994, following the year-end closing of the accounts.

[35]These portfolios would include also those held by the London-based securities firms reported elsewhere in the table. The figures, therefore, cannot simply be added.

Table 2
Selected indicators of leverage in international bond markets

	1991	1992	1993	1994			
				Q1	Q2	Q3	Q4
	in billions of US dollars						
United States	131	99	76	9	–26	–17	–22
Commercial banks[1]	111	105	73	17	– 6	–20	–18
Securities dealers[1]	20	–6	3	– 8	–20	3	– 4
United Kingdom	19	53	136	–43	–18	0	..
Banks:[2] gilts	–2	6	16	2	0	– 1	3
foreign bonds	15	24	52	– 5	– 1	7	19
GEMMs:[3] gilts	..	..	9	– 9	0	– 1	..
Securities dealers: foreign bonds	6	23	59	–31	–17	– 5	3
Total	150	152	212	–34	–44	–17	..
Memorandum items:							
Interbank financed[4]	7	54	182	–54	–48	– 1	17
Repo financed:[5] *Spain*	..	8	24	– 8	– 8	– 4	– 2
Sweden	..	..	13	– 5	– 3	– 6	2

[1] Treasury and agency securities for banks and also including corporate and foreign bonds for securities dealers. [2] Including building societies. [3] Gilt-edged market-makers. [4] Cross-border interbank domestic currency lending by banks in Europe as an indicator of movements in non-residents' bond purchases hedged against exchange rate risk. [5] Indicators of Treasury bond purchases by non-residents financed through repos.

Sources: National data and BIS.

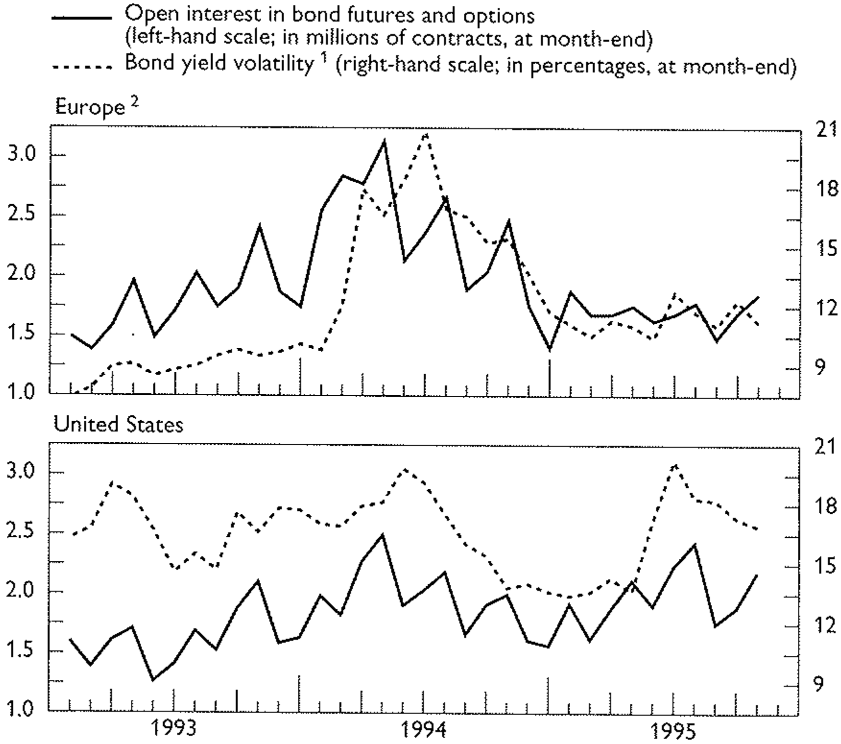
reverse in the first two quarters of 1994.[36] The memorandum items show the same time profile.

The expansion and contraction of positions in bond futures reinforces the impression that leveraged positions bulked larger in 1993 only to be liquidated in early 1994 (Graph 6). Open interest on futures measures the aggregate size of the side-bets on bond prices made on various exchanges. There are two parties to each bet, of course, and thus no net positions. Nevertheless, leveraged investors take long positions, while local arbitrageurs match short futures positions with long cash holdings. Thus, to

[36]Of course, these cash market holdings might well be hedged through a variety of means. On the other hand, these data do not cover the portfolios of other investors that were in effect buying bonds with short-term money, including not only hedge funds but also corporate and municipal treasurers.

Graph 6

Bond futures and options positions and bond yield volatility



¹ As defined in Graph 2. ² For open interest, Belgium, Denmark, France, Germany, Italy, the Netherlands, Spain and the United Kingdom; for bond yield volatility, Germany only.

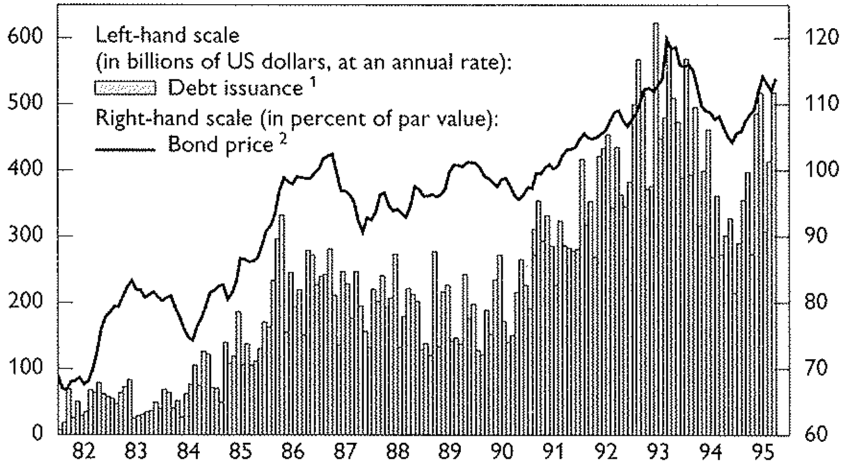
Sources: Futures Industry Association, J. P. Morgan and own calculations.

some extent, open interest can be taken as an indicator of leverage. Open interest peaked in April 1994 in Europe and the United States (Graph 6). In early 1994, as leveraged investors dumped futures, futures prices tended to become cheap relative to cash prices of bonds, and locals would be induced to buy futures and sell cash bonds. In this way, the liquidation of leveraged futures positions would produce selling in the cash market.[37]

[37]See Bensaid and Boutillier (1994).

Graph 7

Bond issuance by US corporations



[1] Gross public debt. [2] Refers to a hypothetical 15-year Baa bond with a ten percent coupon.
Sources: Moody's and the Board of Governors of the Federal Reserve System.

Like leverage, market makers' *management of options* exposures can make volatility directional.[38] The role of portfolio insurance in the volatility of the equity market in 1987 makes one think immediately of option desks hedging puts by selling into a falling market. But according to market participants option trading strategies also played a role in last year's turbulence in bond markets. In their view, options written on behalf of bullish investors in 1993 added to the selling pressure in 1994. Investors who took a leveraged bull position in bonds through options in the later stages of the 1993 bond rally left market makers with exposures

[38]See Bank for International Settlements (1994a), pp. 19–21. Hedging of options imbedded in mortgage-backed securities may produce a symmetric relationship between prolonged market movements and volatility. That is, rising bond yields in 1994 extended the duration of mortgage-backed securities (as refinancing rates slowed). As a result, securities firms and banks needed to short more Treasury notes and bonds to hedge their portfolios. Conversely, falling bond yields in spring 1995 contracted mortgage securities' durations and dealers needed to short fewer Treasury bonds. See Fernald, Keane and Mosser (1994). Consistent with this interpretation and the unique size of the (callable) mortgage market in the US, implied volatility there rose as high in the rally of 1995 as it had in the sell-off of 1994.

that required them to sell as the market fell back into the neighbourhood of the original strike prices.[39]

Asymmetric responses by issuers. Finally, the behaviour of debtors can also contribute to the observed asymmetry. If bond issuers issue into rallies, the increased supply will tend to moderate demand pressures. But if they don't buy in weak markets, they will fail to provide a symmetric support as demand pressures abate. Such *opportunistic issuance* may cut off the peaks of volatility on the upside but fail to do the same on the downside.

The evidence, though sparse, is qualitatively consistent with this picture. Certainly, US corporations issue into rallies (Graph 7). More to the point, governments may do so too. The Bank of England, for instance, acting on behalf of the UK Treasury, dribbles gilt-edged securities into rising markets.[40] And neither corporations nor governments buy in weak markets.[41, 42]

[39]More specifically, market makers who sold call options in the latter stages of the bull market had to purchase the underlying bonds (or futures) as prices rose. As bond prices fell, however, these deep-in-the-money call options again became at-the-money. As the market retraced its steps, market-makers eventually had to sell bonds into a falling market.

[40]Matthews (1995), p. 29, concludes a study of UK government bond issuance of fairly small amounts – 100 to 700 million pounds – of previously auctioned bonds: "the results for conventional tap issues are consistent with the authorities timing conventional taps so as to issue into sectors which are outperforming the rest of the market." This finding is consistent with, but somewhat different from, the policy described in the Bank of England (1993), p. 23: "The principle guiding sales is that the Bank will not sell tap stock into a falling market, but will typically sell successive blocks of stock at progressively higher prices when the price of the stock is rising."

[41] One gilt market participant claimed in an interview that during the period of heavy privatisations in 1986–87, when the public sector borrowing requirement turned negative, opportunistic repurchases led to particularly low volatility.

[42]This failure to repurchase by governments, when contrasted with auction techniques, may lead to an asymmetric effect of large purchases and sales and reinforce some of the mechanisms discussed above. Buying pressure can focus on official auctions and the price discovery process can be diluted over a number of trading sessions through when-issued trading. By contrast, a large, possibly distressed, sale of bonds cannot just wait for such an occasion.

III. Domestic macroeconomic factors

It is now time to consider the link between bond volatility and domestic macroeconomic factors. Our empirical evidence indicates that over long horizons and across countries there is a positive relationship between the level of yields and yield volatility. Since in the long term inflation performance and expectations are probably the main influence on yields, inflation and hence also monetary policy can help to set the background level of volatility. A similar claim can be made for fiscal policy to the extent that it, too, contributes to the level of yields. In the short term, by contrast, the link between such fundamental economic factors and movements in long-term rates cannot be taken for granted. And we fail to find much of a relationship between indices of the volatility in market participants' expectations about inflation or growth and bond yield volatility over the last two years. We do, however, find that apprehensions about fiscal policy probably played some role in Italy.

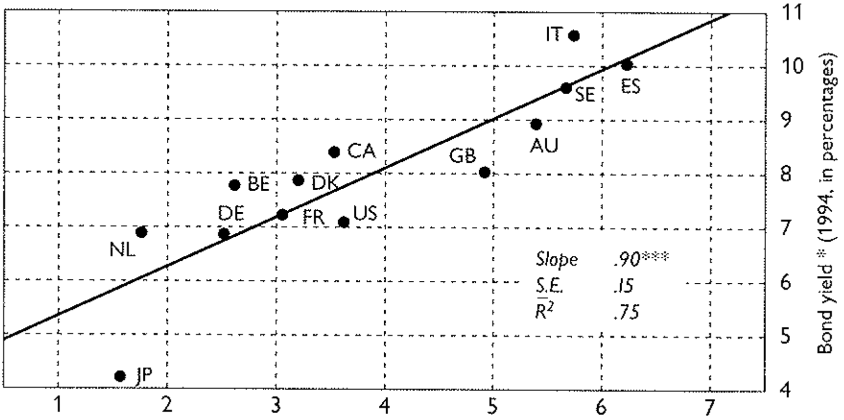
1. *Inflation performance and expectations: the long-term link*

It is widely recognised that, in the long term and in the cross-section, variations in inflation are likely to be the dominant influence on bond rates.[43] The close association between the long-term inflation record and the level of bond yields across countries supports the existence of this relationship (Graph 8). At the same time, there is considerable evidence that bond yields play an important role in setting the background level of volatility, pointing to a role for inflation. Consider the picture that emerges from the US record and from the international cross-section.

[43] The reason is straightforward. One can conceptually decompose long-term rates into two components: expectations of inflation over the relevant horizon, and a residual expected "real" interest rate, including possibly a risk premium. In the long term, it is reasonable to expect that the "real" rate will not diverge too much from the (marginal) return on capital. And over the same horizon, it is equally reasonable to anticipate that the return on capital exhibits a smaller range of variation than inflation, witness the large swings in inflation since the 1970s. Likewise, at a given point in time, arbitrage between bonds denominated in different currencies ensures that the spread between the yields is close to the expected rate of depreciation over the residual maturity of the instruments (plus any, comparatively small, risk premium). Over long horizons, owing to arbitrage in the goods market (international trade) this rate will be largely influenced by the difference between the corresponding inflation rates ("purchasing power parity").

Graph 8

Bond yields and the inflation record



* As defined in Graph 1.

Sources: Datastream and national data.

Bond yields and volatility: the US evidence

In the US bond market, economic historians have established the perhaps surprising fact that the most volatile period in the 130 years since the Civil War coincided with the spell of record-high interest rates fifteen years ago.[44] Their findings are confirmed by the 33 years of daily data at our disposal. When a moving average of monthly yield volatility of the 10-year benchmark bond is plotted against the yield, it is evident that they both peaked in the early part of the last decade (Graph 9). Bond market volatility in the United Kingdom also peaked in the period of high and volatile inflation.[45]

A finer breakdown that juxtaposes US daily yield movements to the level of the yield for the 1980s confirms the association of high rates and high volatility (Table 3). When yields ranged between 7% and 8%, the average absolute change in yield was 4.9 basis points; when they exceeded

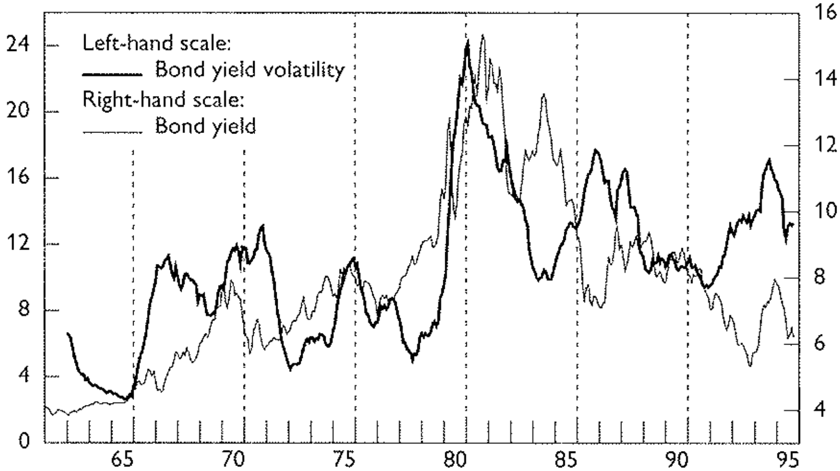
[44]See Wilson, Sylla, and Jones (1990). Though their measure of volatility, which is based on monthly data, is idiosyncratic, it resembles the monthly standard deviation of daily log changes used elsewhere.

[45]See Anderson and Breedon (1996).

Graph 9

Volatility and the ten-year Treasury bond yield in the United States *

In percentages



* Volatility is measured as the twelve-month moving average of the annualised standard deviation of daily percentage changes during calendar months.

Source: National data.

Table 3

Yield and yield change for US 30-year Treasury bond

1st June 1979 to 9th March 1989

Yield range (in percentages)	Number of observations	Average yield (in percentages)	Average absolute change in yield (in basis points)
7–8	269	7.51	4.9
8–9	322	8.75	5.1
9–10	342	9.29	5.3
10–11	423	10.51	6.6
11–12	409	11.56	7.5
12–13	286	12.47	9.8
13–14	286	13.45	10.4
14+	91	14.44	12.5

Source: Ray (1993), p. 458.

14%, it was 12.5 basis points. Since these basis point changes rise faster than the average yield, yield volatility also increases with the level of the yields.

Bond yields and volatility: the cross-sectional evidence

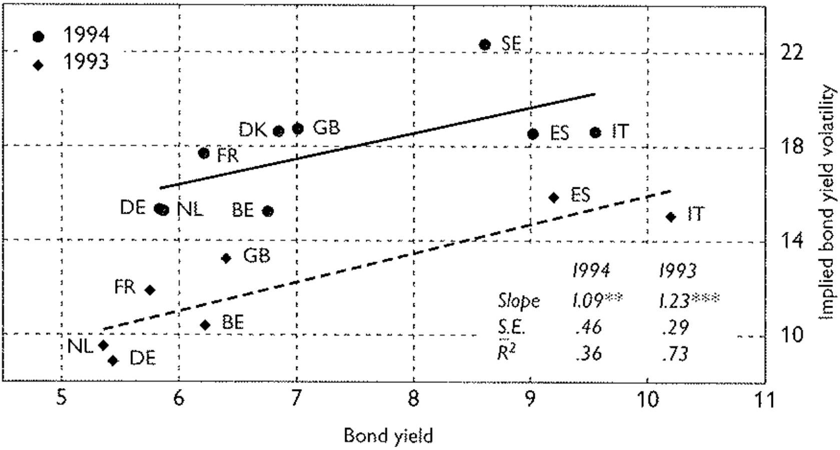
Within Europe, countries with lower yields enjoy less volatile bond markets (Graph 10). In 1993, for instance, the excess of yield volatility of Italian government bonds over that of their German counterparts more or less matched the 4 to 5 percentage point excess of Italian over German government bond yields.

The power of yields to explain differences in baseline levels of bond volatility weakens somewhat when one looks across the G3 bond markets, however. US bond market volatility seems high for Europe. The US combination of yield and volatility would put the US bond market about on the 1994 line for Europe in Graph 10, but leaves it well above in the more normal year of 1993. In fact, in recent years US bonds have tended to exhibit about the same volatility as UK bonds, but a lower yield based on a better inflation record. Even more anomalous is the high

Graph 10

Implied bond yield volatility and yields in European bond markets

Averages of weekly data, in percentages



Sources: J. P. Morgan, Datastream and national data.

Table 4

Volatility of nominal effective exchange rates[1]

	1983[2]–94	1993–94		1983[2]–94	1993–94
	annual averages, in percentages			annual averages, in percentages	
United States	4.9	4.2	Belgium	2.4	3.0
Japan	6.2	7.2	Netherlands	2.8	2.6
Germany	3.3	3.7	Spain	3.5	4.6
France	3.0	2.8	Denmark	3.0	3.2
United Kingdom	5.4	4.9	Sweden	3.7	6.8
Italy	3.3	5.4	Australia	7.5	8.0
Canada	3.4	4.4			

[1] Volatility is measured as the annualised standard deviation of daily percentage changes in nominal effective exchange rates calculated over a calendar month. Nominal effective exchange rates are based on trade flows in manufactured goods between 25 countries.
[2] October–December.

Source: BIS.

volatility in Japan. At the time of writing, Japan is the only industrial country whose bond market shows a yield volatility of over 20%, though yields are comparatively very low by international standards.[46] One clue to explain this anomaly may be the high volatility in Japan's effective exchange rate (Table 4).

Interpretation

The link between bond yield volatility and inflation is consistent with the view that, as suggested by Friedman, inflation becomes harder to predict at higher levels. There is considerable evidence backing this claim. In the US time series, for instance, a review of the record of forecasts of the GDP deflator in the 1970s found the biggest errors in the high inflation years of 1973–75 and 1978–79.[47] Cross-sectional studies and panel

[46] To be sure, the price volatility of the benchmark Japanese government bond is only a little higher than that of German or Dutch bonds. But with a coupon of 4.6% instead of about 7% on the German or Dutch bonds, similar price volatility translates into much higher yield volatility. A recent study has noted the similarity of US and Japanese 10-year bond volatility in the 1990–92 period, but has failed to recognise it as puzzling despite the 2% difference in trend inflation in favour of Japan. See Kikugawa and Singleton (1994), p. 10.

[47] See McNees (1979), p. 50. Khan (1977) revisits Cagan's data with the insight that people incorporate more of an inflation surprise into their expectations as inflation gets higher.

data point in the same direction.[48] This evidence also indicates that the relationship may require considerable dispersion of yields to assert itself. For example, Okun's cross-sectional findings showed that inflation volatility rose only about as fast as average inflation until the latter reached around 4%.

2. *Revisions in, and dispersion of, inflation and growth expectations: a short-term link?*

Over short-term horizons, the link between movements in bond yields and fundamental economic forces such as expectations about inflation is not as well established as over the long-term or across countries at a point in time. For instance, debate continues about whether the change in the level of yields over the last couple of years is explicable in terms of such fundamentals. In general, the shorter the time horizon, the more likely it is that market participants generate yield movements themselves.

In order to test the short-term link with fundamentals, we need measures of uncertainty regarding these fundamentals that are independent of the market movements to be explained.[49] To measure market participants uncertainty regarding inflation and growth, we thus relied on market surveys of the forecasts made by major banks and dealers in the various countries, available at monthly frequencies.[50] Admittedly, this type of data is far from perfect. It is not clear, for instance, whether all respondents update their forecasts monthly, and whether some are not leaning on the consensus forecast. Nevertheless, these forward-looking measures seem more relevant than more mechanical alternatives based on the time-series modelling of inflation and output.

Since the forecasts relate to a horizon of at most two years, the figures for inflation expectations capture only short-term prospects. For ten-year bonds a longer horizon is relevant. Given a typical lag of about

[48]See Okun (1971); Barro (1995); Gruen (1996).

[49]There is a vast, related literature, mostly concerned with equities and bonds, that attempts to establish whether the volatility in asset prices is justified by the *ex post* volatility in "fundamentals" (see Shiller (1979) and (1981) and, for surveys, Scott (1991) and Kupiec (1993)). In contrast to our analysis, this literature does not attempt to explain *changes* in volatility over time on the basis of changes in *observable* macroeconomic fundamentals. The work here is more in the spirit of Schwert's (1989), who relates the (conditional) volatility of stock and (corporate) bonds to the (conditional) volatility of macroeconomic and other factors. His findings point to a limited role for macroeconomic fundamentals.

[50]These were forecasts provided by The Economist and Consensus Economics (1989–1995).

two years between output growth and subsequent inflation,[51] expectations about output growth can provide a useful additional piece of information.[52]

Our evidence on balance indicates that these fundamental macro economic factors do not seem to account for the short-term behaviour of volatility, at least when gauged by the experience of the last few years. This finding appears to hold both on the basis of the frequency and intensity of *revisions* in growth and inflation expectations and of indices of the *dispersion* of such expectations. If an effect exists, therefore, it operates primarily indirectly, notably via induced changes in the *level* of the yields. We do not test for this possibility explicitly. But independent analysis suggests that even the change in the level of the yields over the last couple of years is not easily explicable in terms of fundamentals.[53]

Revisions in growth and inflation expectations

On a priori grounds, one would expect that bond yield volatility would rise as revisions in market expectations about inflation and output growth become more frequent and larger. Such revisions can be taken as a sign of the arrival of new information and/or of greater uncertainty about the future. For instance, the major upward revision in growth expectations in the early part of 1994 for Germany may have partly lain behind the increase in bond yield volatility in that market (BIS *Annual Report* (1995), p. 103).

Nevertheless, a first look at the relationship between the variability of expectations and changes in volatility between 1993 and 1994 hardly points to a significant role of fundamental factors. Despite the sharp rise in the average level of volatility in 1994, the standard deviation of revisions in output growth forecasts actually declined in most countries in 1994; that of inflation forecasts shows no clear tendency (Table 5). Nor can the relative movements in variability explain the cross-country pattern of changes in bond volatility (Graph 11).

This general picture is confirmed by more formal econometric techniques. Even allowing for potential asymmetries in the reaction of implied

[51] See BIS *Annual Report* (1995), Chapter II.
[52] In addition, of course, real interest rates may move with the business cycle.
[53] See BIS *Annual Report* (1995), Chapter V.

Table 5
Volatility of market participants' growth and inflation forecasts[1]

	Growth			Inflation		
	1993[2]	1994[2]	change	1993[2]	1994[2]	change
	in percentage points					
United States . . .	0.11	0.10	–0.02	0.08	0.03	–0.05
Japan	0.25	0.07	–0.17	0.06	0.06	–0.01
Germany	0.17	0.17	0.00	0.04	0.05	0.01
France	0.16	0.06	–0.11	0.10	0.06	–0.04
United Kingdom .	0.06	0.05	–0.01	0.08	0.15	0.07
Italy	0.09	0.14	0.05	0.10	0.06	–0.04
Canada	0.06	0.07	0.00	0.06	0.16	0.09
Belgium	0.15	0.07	–0.08	0.07	0.06	–0.01
Netherlands . . .	0.12	0.11	0.00	0.07	0.08	0.01
Spain	0.10	0.07	–0.03	0.12	0.08	–0.05
Sweden	0.10	0.10	0.00	0.08	0.13	0.05
Australia	0.16	0.12	–0.04	0.09	0.10	0.01

[1] Standard deviation of the monthly changes in the forecast for average annual GDP growth and consumer price inflation respectively over two years. [2] Year in which forecasts are made.
Sources: © The Economist, London (various issues), and own elaboration.

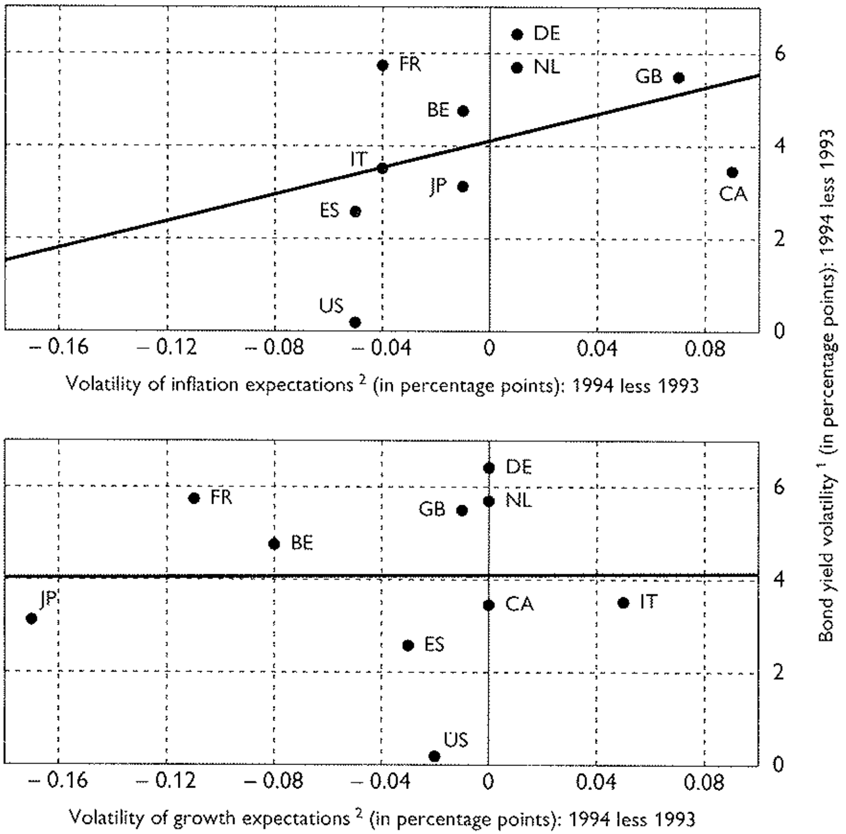
bond yield volatility to revisions in output growth and inflation forecasts,[54] it is very hard to find any evidence of a correct and statistically significant relationship (Table 6). As the few instances shaded in Table 6 indicate, only in a handful of cases are the corresponding parameters consistent with theoretical priors and confidently different from zero, regardless of whether month-end or month average data for volatility are used.[55]

[54]Since the forecasts are made in different months for the current and following calendar year, some adjustment is necessary in order to take care of the varying horizon. We took an arithmetic weighted average of the revisions of the forecasts for the two years. The weight for the current year was set equal to the remaining number of months to its end divided by twelve; the weight for the second year was set at one less the weight for the first year. This index was then added to the standard autoregressions for implied bond volatility, *excluding* the impact of proximate market movements. The latter was excluded because the change in yields could easily have resulted from such revisions in expectations. This potential effect of fundamentals should not be filtered out.

[55]Indeed, the number of statistically significant parameters with the wrong sign is almost as large as that of the correct instances found. Almost as often, for instance, larger revisions in expectations or deteriorating prospects for inflation actually reduce volatility.

Graph 11

Implied bond yield volatility and revisions in inflation and growth forecasts



[1] As defined in Graph 2. [2] Standard deviation of the monthly changes in the forecast for average consumer price inflation and GDP growth respectively over two years.

Sources: J. P. Morgan and The Economist, London.

Dispersion of expectations

To the extent that the *dispersion* of market participants' forecasts at any given point in time is a useful index of uncertainty about the future, one would expect dispersion of inflation and output growth expectations

Table 6

Volatility and revisions in growth and inflation forecasts: regression results[1]

	Month-end				Month-average				Sample period
	ΔInfl⁺	|ΔInfl|	Δgr⁺	|Δgr|	ΔInfl⁺	|ΔInfl|	Δgr⁺	|Δgr|	
JP								18.52*** (5.73)	3:93–5:95
DE			[2]						3:93–5:95
FR	−12.40** (5.85)		20.35** (8.42)		−9.11* (4.96)		13.35** (2.05)		3:93–5:95
NL					12.30** (5.63)				3:93–5:95
SE						17.91** (6.01)			3:94–5:95
AU		−19.12** (7.09)				−13.69** (5.31)			4:94–5:95

[1] Parameter estimates (standard errors in brackets) of the corresponding indicators of revisions in inflation (ΔInfl) and output growth (Δgr) expectations in AR(1) regressions for implied bond yield volatility. See the text for details on the indicators. The table reports only those parameters found to be statistically significantly different from zero. The symbols + and |··| stand for positive revisions and the absolute value of revisions respectively. [2] The parameter for positive growth revisions in Germany was almost statistically significant (11% probability) and equal to 4.72.

Sources: © The Economist and own elaboration.

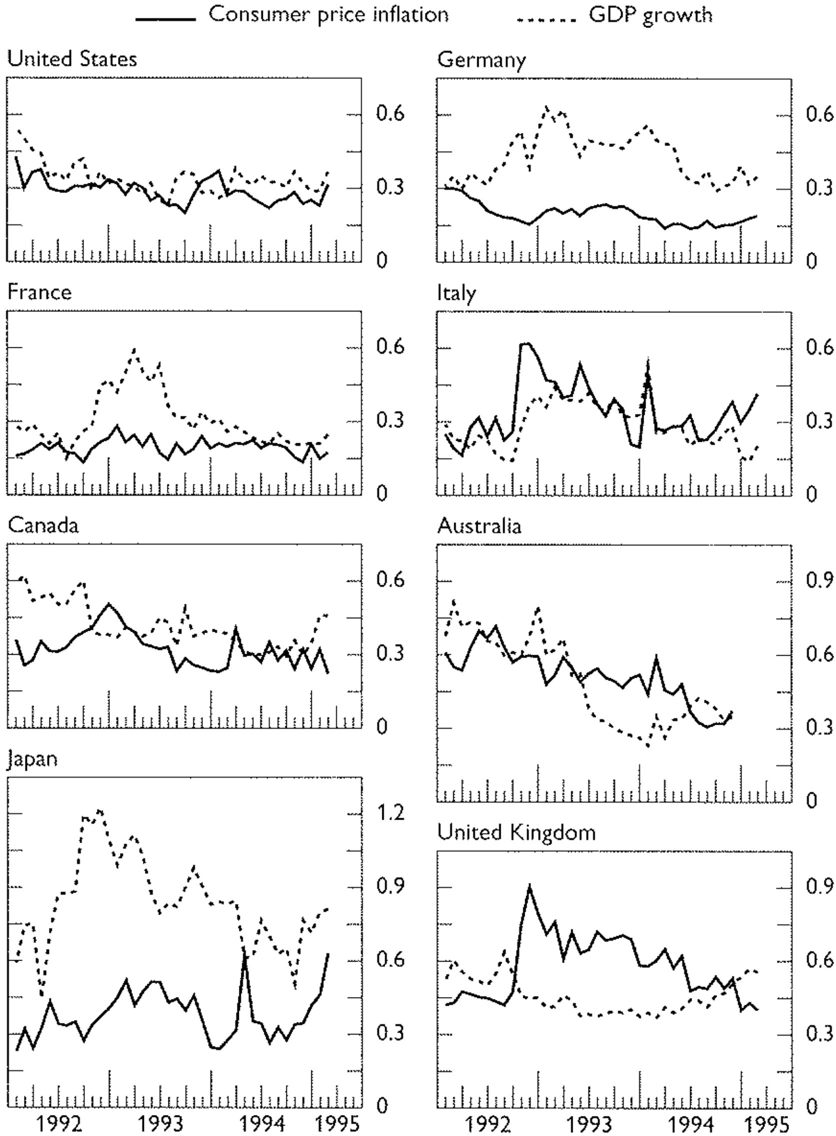
to be positively related to bond yield volatility.[56] A natural measure of such dispersion is the cross-sectional standard deviation of the forecasts made each month by the individual participants.[57]

A first informal look at the evidence provides only limited support for the hypothesis. Admittedly, the index of the dispersion of forecasts of near-term inflation reached a local peak in several countries in the early months of 1994, around the time when bond yield volatility was rising

[56]For a possible formal rationalisation of the link between the dispersion of beliefs and volatility, see Schalen (1993). Frankel and Froot (1990) find evidence of this link in the case of the foreign exchange market at a weekly frequency.

[57]The index was adjusted for the varying horizon in a way analogous to that employed for the correction of the revisions in expectations: we took a weighted average of the standard deviations of the forecasts for the current and following year. This procedure is also used when adjusting implied volatilities derived from exchange-traded instruments for varying maturity effects (see e.g. Canina and Figlewski (1993)).

Graph 12
Dispersion of inflation and growth forecasts*



* Weighted average of standard deviations of forecasts made for the current and following year.
Source: Consensus Economics.

Table 7

Volatility and the dispersion of growth and inflation forecasts: regression results[1]

	Month-end				Month-average			
	Inflation		Growth		Inflation		Growth	
	level[2]	change[3]	level[2]	change[3]	level[2]	change[3]	level[2]	change[3]
JP							3.53* (1.80)	
DE						–33.15* (16.91)		
UK			–9.16** (3.44)				–9.85*** (3.31)	
IT	–8.71*** (3.07)			–5.10* (2.80)				
CA				–5.87** (2.75)				
AU		–56.26** (18.36)						

[1] Parameter estimates (standard errors in brackets) of the index of dispersion of inflation and output growth forecasts, measured by the standard deviation across individual forecasters, in an AR(1) regression for implied bond yield volatility. See the text for details on the index. The table reports only those parameters that are statistically significantly different from zero. No data were available for Belgium, the Netherlands, Spain, Denmark and Sweden. [2] Dispersion at the time the forecast is made. [3] First difference in the dispersion indicator.

Sources: © The Economist and own elaboration.

sharply (Graph 12). And in Japan it increased again in early 1995, in sympathy with the rise in volatility. On balance, however, the dispersion of forecasts of output growth and inflation actually declined between 1993 and 1994, in stark contrast with the global rise in volatility (same graph). Nor can the relative intensity of the change explain the differential increase in volatility across countries (not shown).

The conclusions are even less supportive when the hypothesis is examined through econometric techniques (Table 7). The dispersion index, whether in level or first difference form, is almost invariably statistically insignificant in the standard autoregressions for implied bond yield volatility.[58] And in the few cases where it is confidently different from

[58] As in the case of revisions in expectations, and for analogous reasons, the effect of proximate market movements was not included in the autoregression.

zero, higher dispersion almost invariably is associated with lower volatility.

3. *Fiscal policy*

Just as with inflationary expectations, one can distinguish between the potential role of fiscal policy in explaining the background level of bond yield volatility and its short-term movements. We find some, mainly indirect, evidence that fiscal policy can help to account for cross-country differences in volatility. In the case of Italy, it also appears to explain a small portion of its short-term movements in recent years.

Fiscal policy and volatility: the long-term link

Fiscal policy can contribute to setting the base level of government bond yields. The channel is twofold. In the long term, a weak state of the public sector finances can constrain the freedom of the monetary authorities to pursue a firm anti-inflation policy.[59] Alternatively, and especially in extreme situations, it may give rise to market fears of potential defaults on the debt. And by analogy with inflation, a higher risk premium associated with potential credit risk could also go hand in hand with greater uncertainty about the precise resolution of the fiscal imbalance.

The cross-country evidence supports the link between fiscal policy and bond yields (Graph 13). In a sample of seventeen industrialised countries there is a clear positive relationship between the size of fiscal deficits and the level of bond yields (top panel). This translates into a somewhat weaker relationship with volatility itself (mid panel). The very close association between the fiscal position and the long-term inflation record is consistent with the view that fiscal policy matters largely through the signal it provides about long-term inflation prospects (bottom panel).

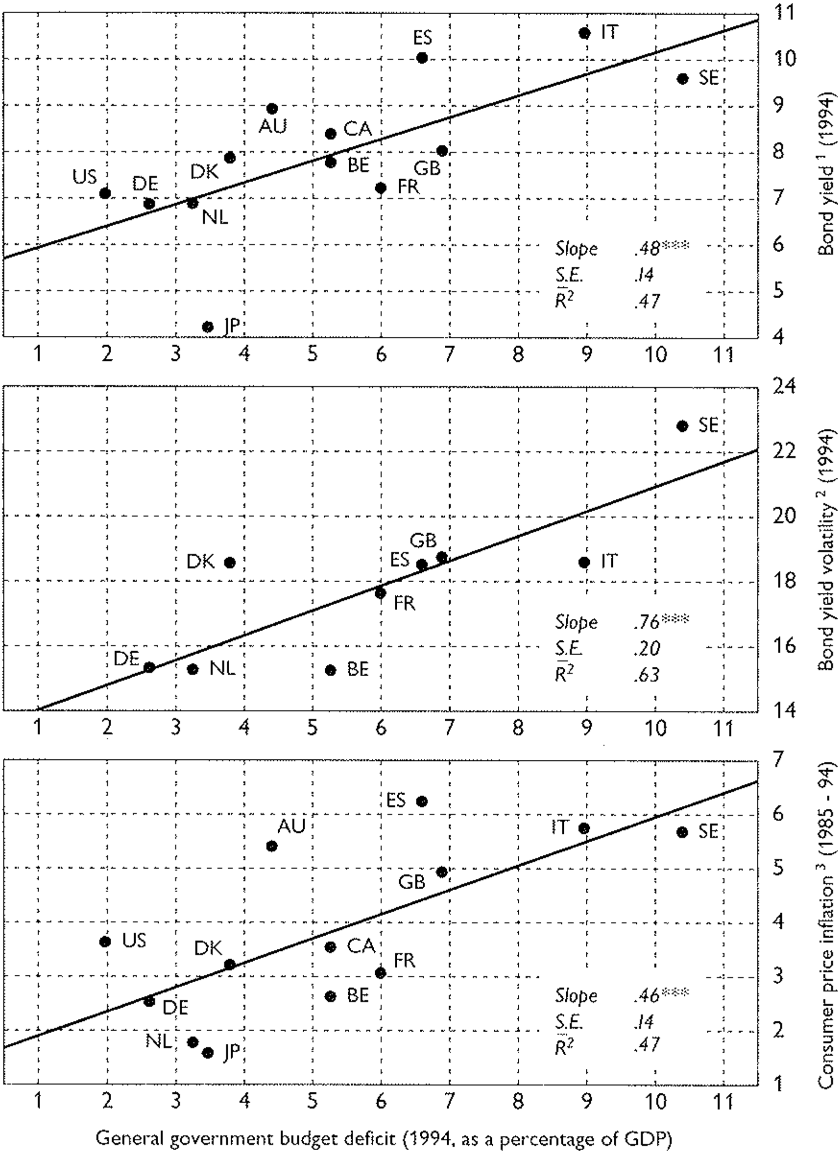
Fiscal policy and volatility: the short-term link

Measuring the impact of changing apprehensions about fiscal policy in the short term is harder. Surveys on short-term budget forecasts are of

[59] Within Europe, given the Maastricht convergence criteria, the budgetary position of a country may also be taken as a barometer of the likelihood and timing of its joining the prospective Monetary Union and hence of shifting to a possibly firmer anti-inflation regime.

Graph 13

Fiscal policy and implied bond yield volatility



[1] As defined in Graph 1 (in %). [2] As defined in Graph 2 (in %). [3] At an annual rate (in %).

Sources: Datastream, J. P. Morgan, OECD and national data.

limited value: the size of the revisions is generally too small, their frequency too low and the forecast horizon too short. We thus concentrated our efforts on the country with the highest ratio of government debt to GDP in our sample, Italy. Given the concerns about the long-run sustainability of the Italian fiscal position, movements in the spread between the government ten-year benchmark yield and the private-sector ten-year swap rate (the "swap spread") can, and often have, been taken as a measure of market participants' changing views of the government's creditworthiness.[60] Indeed, in a configuration unique to Italy, the government bond rate regularly exceeds the ten-year swap rate. In other words, a prime private borrower can borrow lira long-term at a lower rate than the government.[61]

It is in fact comparatively straightforward to point to episodes in the recent past where movements in the swap spread mirrored heightened concerns about the fiscal situation. Thus, a worsening in the perceived credit standing of the government compared with the private sector was a substantial factor in the widening of the swap spread in the summer of 1994 and again in March of 1995 (Graph 14).[62] In the first case, the fiscal background was the delay in the government's passage of a budget through Parliament. In the second, unfolding problems in Mexico were associated with increasing investor intolerance for financing deficits, whether external or internal.

Against this background, we regressed this week's implied bond volatility on last week's implied volatility and on the change over the week in the swap spread. In order to control for broader market movements in long-term rates, we also included the lagged change in the swap rate itself. We permitted widening and narrowing of the spread to exert different influences.

[60] For treatments of the swap spread as a measure of the Italian government's credit standing, see e.g. Banca d'Italia (1995) and Favero, Giavazzi and Spaventa (1995).

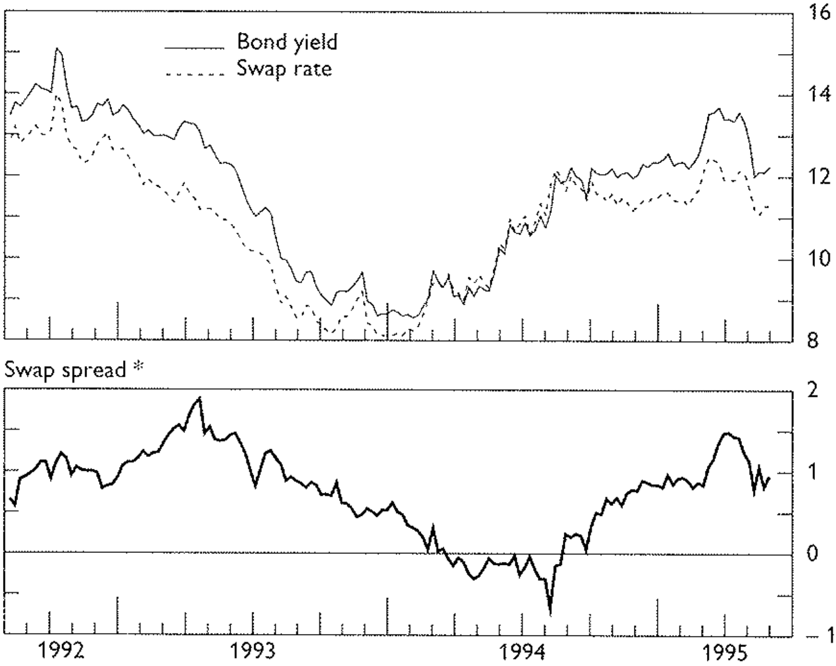
[61] To be sure, there is no consensus on the interpretation of the spread. The configuration is so out of line with experience in other industrialised bond markets that analysts have tended to attribute it to a tax factor, namely the Italian government's slow pace in returning withholding tax on interest payments to foreign residents (Giovanninni and Piga (1992)). After the tax returns were accelerated in April 1994, the private rate did in fact for a time exceed the government rate. However, in the summer of 1994 the government bond rate again rose above the private rate, casting doubt on this interpretation. For a description of the accelerated reimbursement for residents of countries with tax treaties with Italy, see The Treasury (1994), pp. VIII.2–3. More recently, withholding taxes for foreign investors have been eliminated and this policy change, in conjunction with the strong performance of Italian bonds, has narrowed the swap spread. See Banca d'Italia (1996).

[62] Likewise, the government bond yield rose in relation to its German counterpart (Graph 2).

Graph 14

Government bond yield and swap rate in Italy

In percentages



* Difference between the ten-year benchmark government bond yield and ten-year swap rate.

Sources: Datastream and Reuter.

We indeed found evidence of a significant and asymmetric impact. According to our estimates, a ten basis point widening of the swap spread raises implied volatility by a third of a percentage point, while a narrowing of the spread exerts no statistically significant impact.[63] The size of this effect, however, is not very large. The exceptional widening in the autumn of 1994, for instance, at best accounts for a 2 percentage point increase in volatility.

[63]The preferred equation included only positive changes in the swap spread (ΔSP^{+}) and positive percentage changes in the swap rate (ΔRW^{+}, approximated by the first difference in the logs) as controlling variable. Asymmetries are again at work:

$$IVB_t = 2.76^{***} + 2.92^{*}\,\Delta SP^{+} + 0.44^{***}\,\Delta RW^{+} + 0.80^{***}\,IVB_{t-1}$$

(0.65) (1.54) (0.11) (0.04)

IV. Money market volatility

We next examine the link between money-market volatility and bond yield volatility. Just as one can ask what happens to the bond rate when short-term rates change, so one can trace the movements of the volatility of the bond rate associated with changes in money-market volatility. Before reviewing the empirical evidence, however, it is useful to clarify what the link between the two volatilities could mean.

1. *Interpreting the link*

The precise interpretation of the link between money market and bond yield volatility will partly depend on the measure of volatility used. In keeping with our emphasis on uncertainty, we tried to adopt a forward-looking indicator. Unfortunately, we could use implied volatility only for the United States. Otherwise, we relied on the standard deviation of changes in the implicit three-month LIBOR rate three months forward.[64] This rate can be taken as an, admittedly rough, approximation to market expectations of the three-month rate three months hence.[65]

Thinking of money market volatility as indicating uncertainty about the prospects for short-term rates, a natural way of interpreting it is as reflecting uncertainty about monetary policy. While central banks do not *control* money market rates, they influence them closely.[66] Admittedly, the extent to which money market volatility captures uncertainty about policy varies, among other things, with the central banks' willingness to tolerate fluctuations in short-term rates that do not necessarily mirror their policy intentions. But even when such fluctuations are large, the central bank's role is helping to determine money market volatility is in no way diminished.

[64] This is essentially the standard deviation of the rate on forward rate agreements (Bank for International Settlements (1986), Chapter IV).

[65] More precisely, money market volatility is measured by the standard deviation in the daily percentage changes (approximated by the logarithmic first differences) in the implicit three-month LIBOR rate three months forward. The window differs depending on the specific application. At least at the weekly frequency, when the window is only five business days, we measure the standard deviation around an imposed zero mean, a measure known as "quadratic mean". The reason is that it does not seem advisable to let very short-term trends (i.e. the mean percentage change during a week) hold down the measure of variation.

[66] They do so by affecting conditions in the market for bank reserves, notably by setting the terms on which banks can borrow from, or lend to, them. See, e.g., Kneeshaw and Van den Bergh (1989) and Borio and Fritz (1995).

Money market volatility can cause bond yield volatility, but the reverse is also possible. This is true regardless of whether money market volatility captures monetary policy or not. In the case of policy, were a central bank expected to respond to higher yields as an indicator of higher inflation, then movements of yields at the long end might be interpreted as making more likely changes in rates at the short end. Alternatively, and of particular relevance in 1994, a portfolio adjustment by a leveraged bond holder may lead to a similar result. The investor sells off his holdings of bonds and in the process reduces his demand for short-term funds in the repurchase market. Essentially, the disturbances to the two ends of the yield curve are just different legs of the same transaction. Money market observers might see a drop in demand for short-dated funds as the result of repositioning of bond portfolios. The use of causal language in what follows should therefore be interpreted with some caution.[67]

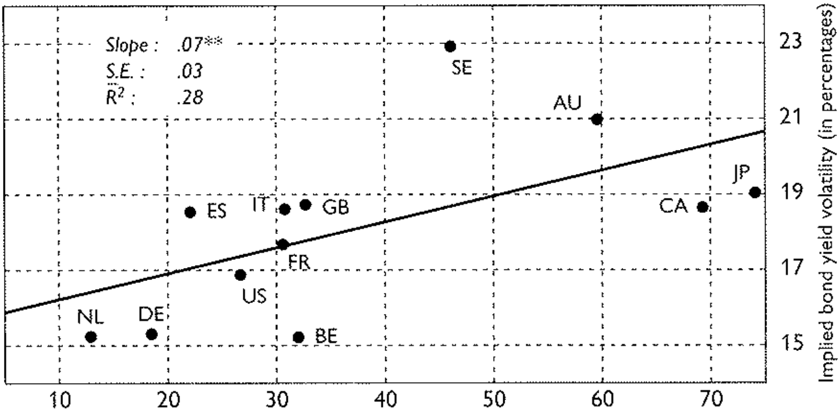
2. *Evidence*

Cross-sectional evidence and the US long-term time series indicate that money market volatility helps to set the background level of bond yield volatility. In the more recent period, formal econometric analysis also points to a link with the short-term movements of bond volatility. To the extent that money market volatility reflects the operation of, or uncertainty about, monetary policy, our evidence points to another channel through which policy can affect volatility. In the time series, the strength of the link is best appreciated when implied, rather than realised, money market volatility is used in the regressions. Given the observed range of variation of money market volatility, the variable appears to contribute

[67]What should the strength of the link be expected to be? The response of long-term rates to short-term rates can serve as a basis for deriving an expectation. If the long rate responds by a given fraction, then the volatility of the long rate should move by the same fraction with respect to the volatility of the short rate. Central bank estimates of the effect of short on long rates over a one-year horizon typically range between 0.2 and 0.4 for most G-10 countries. (Smets (1995), p. 244). Germany, Switzerland and Canada are outliers on the low side and Italy on the high side (but the long rate for Italy has a rather short maturity). Other estimates reviewed in Akhtar (1995), p. 120, for the United States based on daily data suggest that a 1% rise in the Federal funds rate yields a 12 or 13 basis point increase in the ten-year rate over a one-day or ten-day interval. These are smaller effects than the 19 basis points over a year reported by the Board of Governors in Smets (1995). On the other hand, there is some evidence that the pass-through has become stronger in recent years (Cohen and Wenninger (1994) and Estrella and Hardouvelis (1990)). On balance, the 0.2-0.4 range should be viewed as a high upper bound for relationships measured at the higher frequency of one week.

Graph 15

Implied bond yield volatility and money market volatility



* Annualised standard deviation of the daily percentage change in the yield on three-month LIBOR three months forward; monthly average for 1994. The measure avoids the direct influence of the authorities on spot short-term rates and is therefore a better indicator of market expectations.

Sources: J. P. Morgan, national authorities and BIS.

more to the explanation of cross-country differences than to changes over time in bond volatility. In a number of countries, notably the United States, the link with bond yield volatility appears to strengthen in the period following the Fed's tightening of monetary policy in February 1994, but the relationship is not such as to explain much of the increase in bond yield volatility worldwide.

The cross-sectional and longer-term US evidence

In the cross-section, higher money market volatility was associated with higher bond yield volatility across a dozen markets in 1994 (Graph 15). The estimated strength of the relationship, a 5% pass-through, is sufficient to offer an important clue to national differences in bond market volatility. Consider the gap between the core European countries' money market volatility, in the teens, and the high volatility characteristic of, say, the Tokyo money market. To use round numbers take the former to be 20% and the latter to be 60% (Table 8). A 5% pass-through coefficient applied to the 40% difference in money market volatility yields 3

Table 8

Bond and money market volatility in 1993 and 1994

	Bond market[1]				Money market[2]				
	1993	1994	change	peak 1993–94	1993	1994	change	peak 1993–94	1994
	in percentages			date	in percentages			date	
US	16.7	16.9	0.2	June 1994	28.2	26.7	– 1.5	May 1994	
JP	15.9	19.0	3.1	March 1994	57.6	74.1	16.5	Jan. 1994	
DE	8.9	15.3	6.4	June 1994	16.1	18.5	2.4	June 1994	
FR	11.9	17.7	5.8	June 1994	43.0	30.6	–12.4	July 1993	March
UK	13.3	18.7	5.4	June 1994	29.1	32.7	3.6	June 1994	
IT	15.1	18.6	3.5	July 1994	32.3	30.8	– 1.5	Nov. 1993	Sept.
CA	15.2	18.6	3.4	July 1994	52.8	69.3	16.5	Feb. 1994	
BE	10.4	15.2	4.8	July 1994	47.8	32.0	–15.8	July 1994	
NL	9.5	15.2	5.7	July 1994	15.8	12.9	– 2.9	July 1993	May
ES	15.9	18.5	2.6	July 1994	33.3	22.1	–11.2	July 1993	Aug.
SE	–	22.9	–	Sept. 1994[3]	55.1	46.1	– 9.0	Sept. 1993	Aug.
AU	–	21.0	–	April 1994[4]	34.4	59.6	25.2	April 1994	

[1] Implied volatility (see Table 1 for details); averages of weekly data. [2] Money market volatility is measured by the annualised standard deviation of the daily percentage change in the yield on three-month LIBOR three months forward calculated over calendar months; averages of daily data. [3] Peak during February–December 1994. [4] Peak during March–December 1994.

Sources: J.P. Morgan, national authorities and BIS.

percentage points on the bond volatility. This 3% is a sizeable part of the gap between Japanese and German bond volatility (same Table).[68]

In modern US history the highest bond yield volatility coincided with the highest money-market volatility about fifteen years ago (Graph 16).[69] Professor Scylla's observation that one of Paul Volcker's "achievements" was to impart to bonds the volatility usually associated with stocks points to the effect of close targeting of banks' (non-borrowed) reserves on the volatility of short and long interest rates. It may seem that our account of the high bond yield volatility of the early 1980s now suffers from an over-

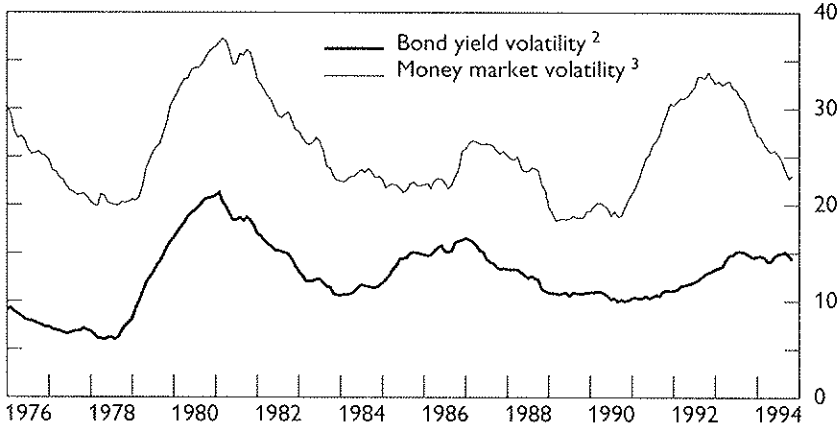
[68]We attempted to distinguish between the relative contributions of money market volatility and the level of bond yields or long-term inflation performance. However, the correlation between those variables in the cross-section was so high (94%) that it was impossible to do so with any degree of confidence. This raises the question of how stable this correlation is and, if stable, what could explain it.

[69]See Wilson, Scylla and Jones (1990).

Graph 16

Bond yield and money market volatility in the United States [1]

In percentages



[1] Centred 25-month moving averages of realised bond and money market volatility, measured as the annualised standard deviation of the daily percentage changes in the corresponding rates, calculated for calendar months. [2] Based on the ten-year benchmark government bond yield. [3] Based on three-month LIBOR three months forward.

Sources: J. P. Morgan, national authorities and BIS.

abundance of explanations: high inflation and high money market volatility. At some level, however, these two explanations are not alternatives. Rather, the Federal Reserve's willingness to see greater fluctuations in short-term rates was itself a response to the unprecedented challenge of high inflation expectations as embodied in bond yields.[70] In other words, the high money market volatility ultimately reflected the difficult position of the central bank in the face of runaway inflation expectations.

The more recent evidence: realised money market volatility

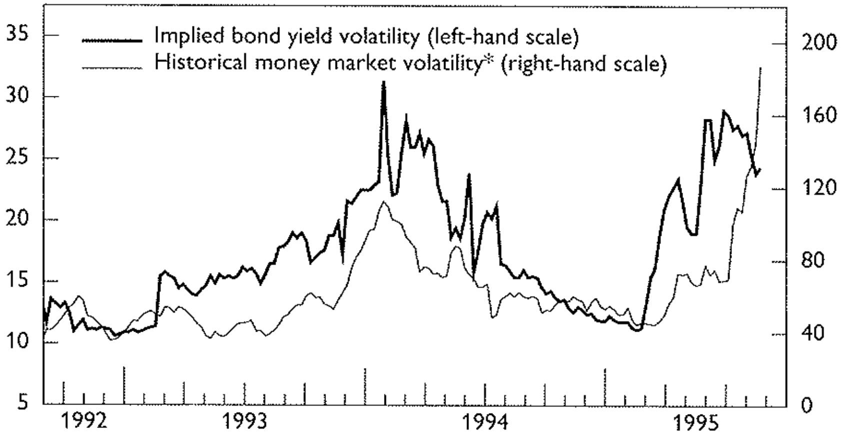
A first informal look at the experience since 1992-93 provides rather mixed results on the role of realised money market volatility. Peaks in money market volatility sometimes coincided or just preceded those in implied bond yield volatility. For the years 1993 and 1994, they coincided in Frankfurt, London and Brussels and led them slightly in New York and

[70] See e.g. Goodfriend (1995).

Graph 17

Implied bond yield volatility and historical money market volatility in Japan

In percentages



* Annualised weekly volatility, calculated over a one-week window, with an imposed zero mean; nine-week moving average.

Sources: J. P. Morgan and national authorities.

Tokyo (Table 8). Moreover, in some countries the existence of a link is obvious to the eye. In Japan, for instance, the rise in bond yield volatility in January 1994 echoed developments in the money market (Graph 17). On the other hand, the strains in the European Monetary System in 1993 rocked money markets in France, the Netherlands and Spain without pushing bond market volatility to peaks (Graph 2). And, on average, money market volatility actually fell in 1994 in a majority of countries despite the rise in bond yield volatility (Table 8).

Econometric evidence at a higher frequency is more supportive of a general link. In this case we added *realised* money market volatility to the standard autoregressions for implied bond yield volatility.[71] We performed the exercise both at the weekly and monthly frequency, with money market volatility being measured over the corresponding window. We carried out the exercise also on monthly data since the measurement

[71] The results controlling also for the impact of proximate market movements are broadly similar but the link appears a bit weaker (Annex V).

Table 9

Implied bond volatility and realised money volatility: regression results[1]

	Weekly			Monthly[2]		
	whole sample	earlier period	later period	whole sample	earlier period	later period
United States . .	0.012**	0.005	0.018**	0.025	–0.018	0.051
	(0.005)	(0.006)	(0.007)	(0.025)	(0.032)	(0.045)
Japan	0.004	0.018	–0.005	0.088**	0.079**	0.040
	(0.007)	(0.011)	(0.008)	(0.037)	(0.032)	(0.040)
Germany	0.025**	0.010	0.032**	0.050	0.038	0.044
	(0.010)	(0.008)	(0.015)	(0.071)	(0.049)	(0.131)
France	0.005	0.004	0.010	0.037	0.027**	0.129[3]
	(0.005)	(0.005)	(0.012)	(0.024)	(0.012)	(0.073)
United Kingdom .	0.009*	0.011*	0.015	0.017	0.028	0.093
	(0.005)	(0.005)	(0.017)	(0.025)	(0.028)	(0.178)
Italy	0.011	0.011	0.017*	0.024***	0.029**	0.041
	(0.010)	(0.015)	(0.009)	(0.009)	(0.010)	(0.050)
Canada	0.004	0.009*	0.001	0.018	0.051***	–0.005
	(0.003)	(0.005)	(0.002)	(0.012)	(0.017)	(0.014)
Belgium	–0.003	0.004	–0.012	0.018*	0.023*	0.026
	(0.006)	(0.003)	(0.010)	(0.010)	(0.011)	(0.019)
Netherlands . . .	0.017***	0.001	0.054***	0.017	0.018	0.062
	(0.006)	(0.004)	(0.017)	(0.023)	(0.014)	(0.111)
Spain	0.006		0.003	–0.048	–0.076	0.006
	(0.006)		(0.010)	(0.035)	(0.044)	(0.029)
Denmark			0.020*			0.125**
			(0.011)			(0.047)
Sweden			0.023*			0.134***
			(0.009)			(0.041)
Australia			0.009			0.096**
			(0.008)			(0.041)
Japan (period split at end-1993)	0.004	0.004	0.003	0.088**	0.044**	0.122*
	(0.007)	(0.007)	(0.010)	(0.037)	(0.019)	(0.059)

[1] The table reports the coefficient of money market volatility in an AR(1) regression for implied bond yield volatility. Money market volatility is measured as the standard deviation (around an imposed zero mean) of the implied three-month LIBOR three months forward calculated over non-overlapping one-week horizons (Friday to Thursday). Standard errors are shown in brackets. [2] Month-end data. [3] Marginal significance level equal to 10.2%.

Sources: J.P. Morgan, national data and own elaboration.

of money market volatility over such a longer window can prove useful: especially in countries with fairly high volatility, the measure is less sensitive to very short-term, possibly exceptional, developments. This benefit, however, comes at the cost of losing observations on bond yield volatility.[72]

For almost all of our countries, regression analysis shows a positive relationship between money market and bond market volatility. For seven out of the thirteen cases, the relationship is evident at the weekly frequency (Table 9). In New York, Frankfurt and London, one or two percent of Friday through Thursday's measured money market volatility shows up in the respective Thursday close implied bond volatilities. A similar "pass-through" is found for Amsterdam, Milan, Copenhagen and Stockholm. The rest of our thirteen markets show no significant relationship at the weekly frequency. These negative findings include all of the money markets showing money volatility above 50% on an annualised basis. Most probably, in these cases the link is disturbed by the sharp but short-lived movements in short-term rates in response to strains in the foreign exchange market.

At the monthly frequency, a positive relationship emerges for five of the centres where it was not evident in the weekly data: Tokyo, Paris, Toronto, Brussels and Sydney. The link at this frequency tends to be a bit stronger. In particular, as much as 9% of money market volatility is transmitted to bond yield volatility in Tokyo.

The relationship between money market and bond yield volatility seems to strengthen in the latter half of our sample period (same Table). In the United States, for instance, in the period to the end of January 1994, before the Fed's first tightening move, the coefficient linking the two was essentially zero; in the period beginning in February 1994, the effect weighed in at almost 2%. A similar pattern is found for Germany and the Netherlands; Britain is one notable exception.

The more recent evidence: implied money market volatility in the United States

On both conceptual and statistical grounds, it would clearly be more appealing to match up *implied*, rather than *realised*, money market

[72] Annex IV discusses these issues in detail and describes an econometric technique designed to improve this trade-off.

volatility with implied bond yield volatility. Conceptually, the two measures are closer. Statistically, some of the measurement problems are avoided.

Owing to data limitations, we could only perform this exercise for the United States. The measure of implied money market volatility is that derived from interest rate caps. These at-the-money volatilities are the key pricing parameter from the over-the-counter market in caps or floors on LIBOR-based assets or liabilities.[73]

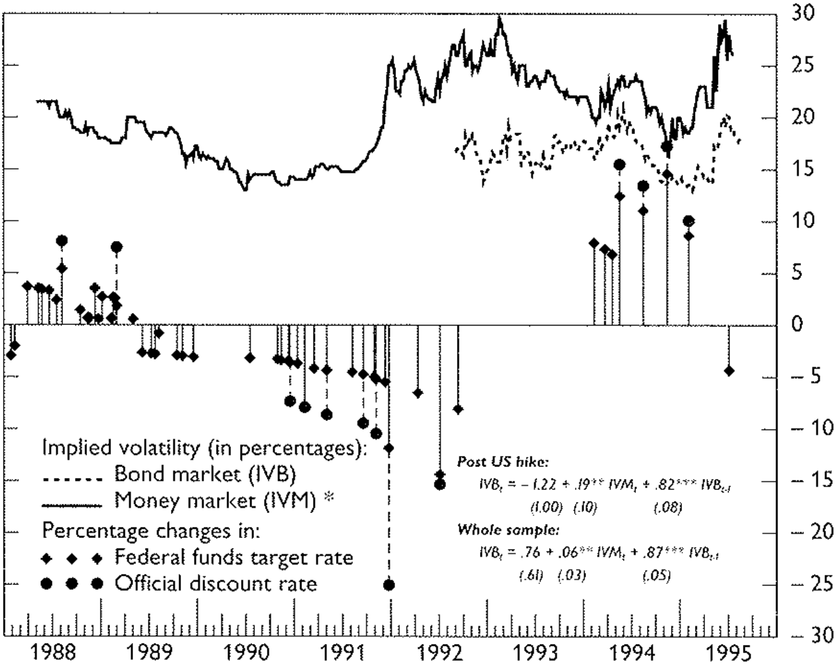
An informal juxtaposition of implied money and bond yield volatility confirms and strengthens the previous evidence derived from realised money market volatility (Graph 18). The two volatilities tracked each other much more closely in 1994 than in the earlier part of the sample period (since August 1992). And the econometric results point to a much more powerful transmission of volatility. Over the whole sample, it is of the order of 5-6%, depending on whether lagged implied money market volatility is also considered (same Graph). This result suggests that our crude weekly measure of the variability of short-term rates can understate the link by a factor of 4 or 5.[74] Similarly, the strengthening of the relationship in the later period stands out even more clearly. Prior to February 1994, no transmission is apparent; thereafter close to 20% of money market volatility shows up in bond yield volatility.

If the point estimate of the pass-through is taken as a benchmark, then money market volatilities could help to explain a considerable portion of the rise in US bond yield volatility in 1994, given that this increase was itself comparatively small by international standards. By contrast, even the upwardly adjusted estimates indicate at best a very modest role elsewhere. Indeed, most European money markets – with Frankfurt the notable exception – were calmer in 1994 than in 1993.

[73]A corporate treasurer who is comfortable paying the current LIBOR of, say, 5% but who fears that rises in LIBOR will make his floating-rate liabilities increasingly costly to service might contract for a cap. For a premium up front, the treasurer's counterparty, say the provider of our data, Chase Manhattan, contracts to pay any excess of interest costs over the current LIBOR of 5%. The data are for 3-year caps. Such a cap can be considered a strip of options, with maturities at 3 or 6-month intervals, extending from the first covered payment to the maturity of the cap, in this case three years. For an analysis of Eurodollar volatility derived from exchange-traded options, see Abken (1995).

[74]This result is confirmed by the use of the alternative econometric technique used to adjust the estimates derived from realised volatilities (Annex IV). That technique also suggests that the adjustment factor may typically be somewhat lower for other countries, generally ranging between 2 and 4.

Graph 18
Implied bond yield and money market volatility and monetary policy in the United States



* Derived from three-year caps on three-month LIBOR.
Sources: Chase Manhattan, J. P. Morgan and the Federal Reserve Board.

As regards the possible factors behind movements in US implied money market volatility, Graph 18 suggests a link with monetary policy. Most clearly, the unprecedented 1% discount rate move in December 1991 and, later, the Fed tightening in February 1994 and again in November of the same year were all associated with increases in money market volatility. Nevertheless, the size of the potential effect should be kept in perspective. Even if, say, the full 6.5 percentage point rise in implied money market volatility in December 1991 was wholly ascribed to the change in the discount rate, the corresponding impact on bond yield volatility would range between less than 0.4 and 1.5 percentage points.

V. International factors: spillovers and (dis)investments

So far, we have considered only the domestic determinants of bond yield volatility. We next examine possible international influences and focus on three issues: contemporaneous correlations between markets; spillovers, in which case lead and lag relationships are duly taken into account; and the impact of international capital flows.

Confirming previous analogous findings for stock markets, mainly concerned with the 1987 Crash, we find evidence that cross-country correlations of volatility were generally higher, and spillovers stronger, during the bond market turbulence of 1994. The hierarchy of influence of spillovers across markets, however, differs somewhat from that typically found in stock market studies. In addition, we show that international disinvestments appear to have rocked European bond markets at the time.

1. *Contemporaneous correlations and spillovers of volatility*

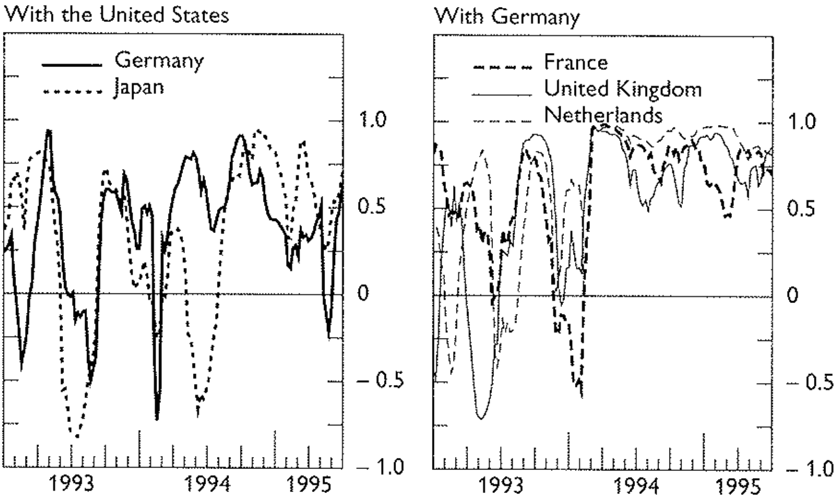
Events in 1994 are consistent with the view that volatility is more correlated across markets when it is high.[75, 76] The rolling estimates shown in Graph 19 indicate that cross-country correlations tended to increase in the period of bond market turbulence in 1994, when bond yield volatility generally rose (Graph 2). Among the G-3, the correlation between volatility in the German and US markets was both higher and more

[75] See Singleton (1994b) for this result for US and Japanese bonds. Studies that find spillovers from one equity market to another in the form of large movements in one market's index raising estimated volatility in another market test for a restricted version of correlation. That is, they test for correlation of volatility that is not the result of correlation of the underlying cash markets.

[76] This view should not be confused with the claim that when volatility is high, cross-market correlations of *price movements* (or *returns*) rise. The truth of this second claim matters to portfolio managers, who hope that low correlations will hold down aggregate risk in their portfolios, and to risk managers of trading books, whose models of value at risk are liable to understate risk if correlations ratchet up in falling markets. The claim, in combination with the directionality of volatility, amounts to a variant of Sod's or Murphy's Law – just when you need the benefit of diversification, it proves least in evidence. Some empirical support for the claim has been produced for the stock market. Correlations of returns rose around the 1987 Crash, according to one study (Bennett and Kelleher (1988)). Its validity for the bond market has been disputed by Singleton, analysing data on US and Japanese bonds (Singleton (1994b)). Other evidence, however, indicates that it seems to hold in parts of the dollar bond market. One study has shown that Brady bonds, securities created out of the bank debt of such countries as Mexico, Argentina and Brazil, performed much more like Treasury bonds in the volatile months of 1994 than they had in 1993 (Clark (1994)). Suffice it to note that the volatility-correlation link in world bond markets represents an important question for further research.

Graph 19

International correlations of implied bond yield volatility *



* The correlation coefficient between weekly implied yield volatilities is calculated over a sixteen-week sliding window and is plotted at the point corresponding to the last observation.
Sources: J. P. Morgan and own calculations.

consistent after February 1994 than before. Within Europe, volatility evolved fairly closely. This was in contrast to developments in the summer of 1993, when ERM strains interrupted the usual co-movements between the volatility in Germany, on the one hand, and in France and the Netherlands, on the other.[77] The exception to the general pattern in 1994, as so often in this analysis, was Japan.

Testing for spillovers

Contemporaneous correlations leave open the question of whether common movements in volatility amount to anything more than common reactions to incoming news and more general shocks, or indeed common subsidence in volatility following such shocks. The econometric literature has addressed this issue by considering leads and lags in the relationships, adding lagged foreign influences to the domestic volatility process and

[77] This was true regardless of the direction of exchange rate pressures. The French franc had come under downward pressure while the Dutch guilder tended to appreciate.

interpreting the resulting links as "spillovers".[78] Here we do no more than follow this tradition, while fully aware of the difficulties in interpretation. The novelty is that, to our knowledge, no published study has as yet examined the bond markets or used implied volatilities.[79] As a first step we consider the effect of last week's implied volatility in, say, the US bond market on this week's implied volatility in the German market, controlling for last week's implied volatility in the German market.[80]

Generally, spillovers assert themselves in periods of high volatility when the persistence (autoregressive) model breaks down. For a number of countries, spillovers add explanatory power over and above domestic lagged volatility precisely when this falters (Graph 20). Since both contemporaneous volatility and these rolling regressions signal different dynamics in 1994's more volatile market, we investigated the global structure of spillovers before and after the break represented by the Federal Reserve's first tightening move of February 1994.

Spillovers vary in extent and direction over time (Graphs 21 and 22). In the earlier period, Frankfurt's implied volatility inflected that in London, Amsterdam and Milan, while London's exerted some influence over that in Brussels.[81]

[78]There is a vast such literature on equity and foreign exchange markets. For studies of equity markets, see Bennett and Kelleher (1988), Roll (1989), Wadhwani (1989), King and Wadhwani (1990), Theodossiou and Lee (1993), Lin, Engle and Ito (1994), Susmel and Engle (1994), Wei, Liu, Yang and Chaung (1995). For studies of foreign exchange markets, see Engle, Ito and Lin (1990), Baillie and Bollerslev (1991), Ito, Engle, and Lin (1992), Baillie, Bollerslev, and Redfearn (1993). The chosen aquatic imagery of spillovers may seem a bit slow for an electronic age, and the speedier image of a meteor shower may serve better (although it may reflect professional envy by economists of the forecasting ability of astronomers). The popular image of the meteor shower suggests volatility travelling from east to west as the globe turns, as traders take their cue from movements in the succesive opening and closing of markets (Engle, Ito and Lin (1990)). These studies make extensive use of daily and intra-day data, a higher frequency than the one we could examine in our case.

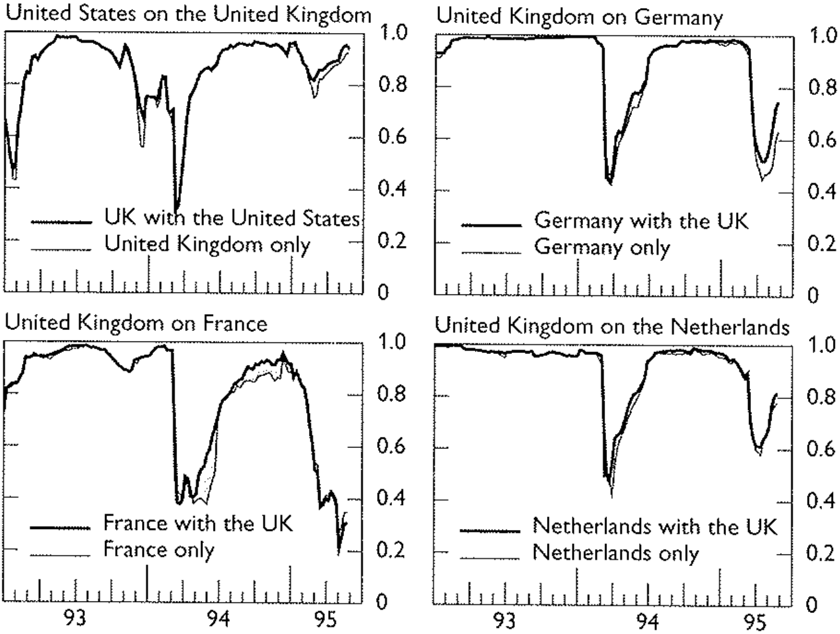
[79]The great advantage of using implied volatility is that no degrees of freedom are used up in estimating the volatilities themselves, a major drawback of GARCH models. We can thus concentrate on finer issues regarding the stability of the process over short horizons. The disadvantage of our data set is that only weekly observations were available.

[80]Here we depart from the approach in the cited studies. They test for a volatility spillover as defined by an effect on market *i*'s (estimated) volatility of a lagged jump in the price in market *j*, given the effect of any lagged jump in the price in market *i*. By contrast, we test for a volatility spillover as defined by an effect on market *i*'s implied volatility of lagged implied volatility in market *j*, given lagged implied volatility in market *i*. For detailed results, see Annex VI.

[81]It may seem odd that German bond market volatility affected volatility more in the bond markets linked by floating than by pegged exchange rates *vis-a-vis* the German mark, that is those in London and Milan rather than in Paris. We hypothesise that the currency strains of 1993, and the reaction of short-term interest rates to them, tended to uncouple volatility across all except the most closely linked core European bond markets of the Netherlands and Germany. This hypothesis is consistent with the previous observation that the contemporaneous correlations in volatility declined at times of strains in the currency markets.

Graph 20

The explanatory power of persistence and spillovers: rolling regressions *



* Uncentred $\bar{R}^2$ from (de-meaned) AR(1) rolling regressions for market i to which the previous week's volatility in market j is added. The regressions are run over a sixteen-week window.

Sources: J. P. Morgan and own calculations.

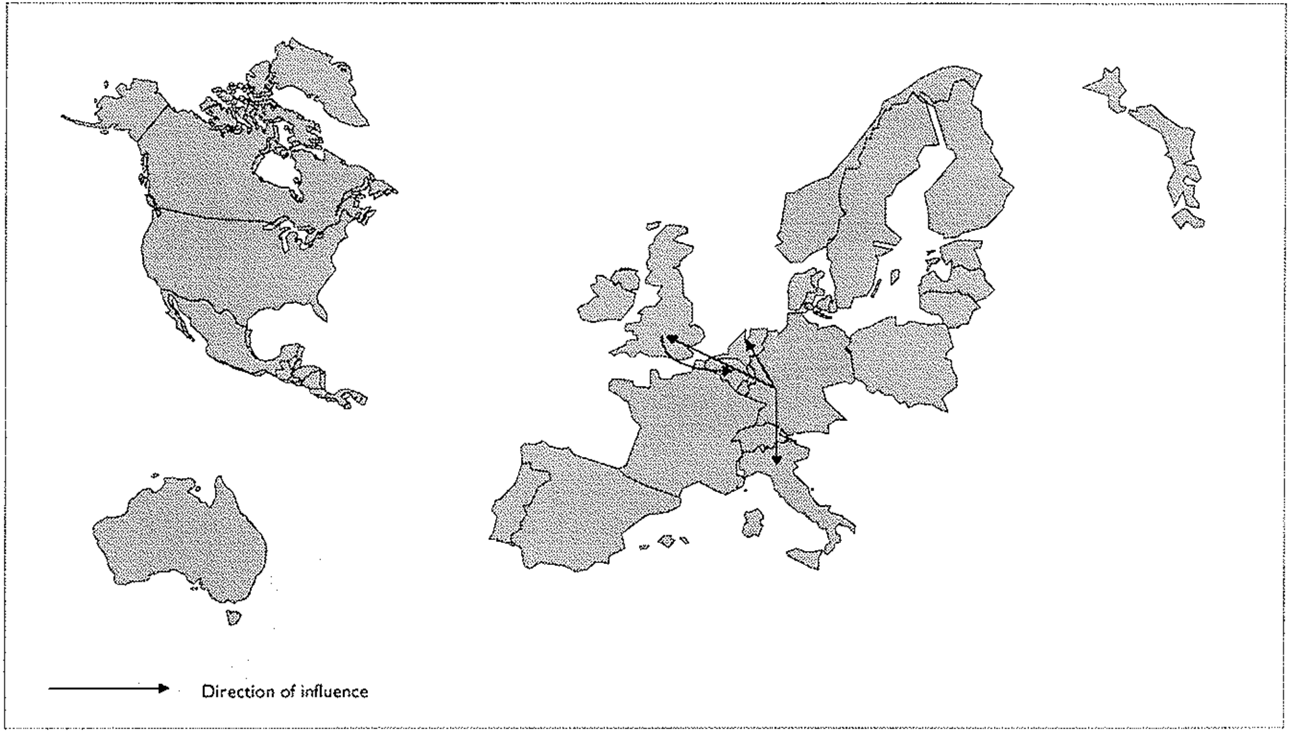
Spillovers became much more pervasive in the later period, and New York and London gained importance as a source of transmission. Only Tokyo remained neither a source nor a recipient of volatility.[82] Volatility in New York made its impact felt in all financial centres except Tokyo. This influence was asymmetric, with none of these centres appearing to affect volatility in New York. The propagation seems to have proceeded along two different paths: a set of "high-yielders" (Sydney, Toronto, Milan, Stockholm and Madrid) and the set of remaining European financial centres (London, Frankfurt, Paris, Amsterdam and Brussels). In the

[82]Tokyo's status as a fixed-income hermit sets our findings apart from those of studies of the transmission of stock market volatility.

Graph 21

Volatility spillovers

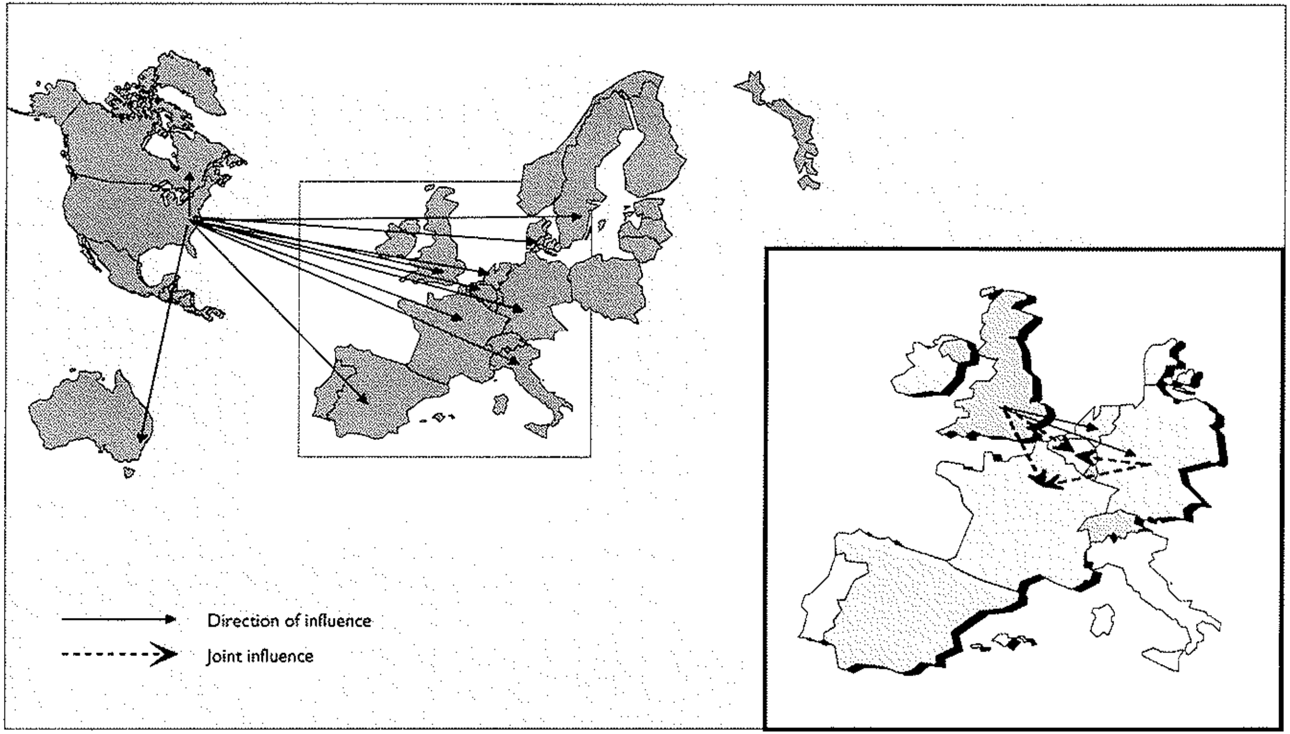
August 1992–January 1994



Graph 22

Volatility spillovers

February 1994–May 1995



second group, London plays a key role, turning the tables on Frankfurt and transmitting its volatility more widely.[83]

Interpreting the pattern of spillovers

The two key findings regarding spillovers have precedents in research on equities. Volatility spills from one market to another with greater force in the period that includes the 1987 stock market Crash.[84] Thus, the formal tests repeat the message of the simpler observation that volatility is more correlated at high levels of volatility.[85] It is as if traders can take their cue from foreign markets' prices only when they are jumping around enough to be seen over the horizon.

The hierarchy of markets emerging from our analysis also has some precedent in the analysis of equity markets, although the details and articulation look quite different. The role of the US market as an unmoved mover is not altogether surprising. By contrast, the somewhat hermetic, or hermit, position of the Japanese bond market is unusual. One is tempted to identify the transmission of US volatility through London to the Continent as somehow reflecting the location of much of the European options market in London, but perhaps one should consider the finding as a symptom of the greater macroeconomic integration between Britain and the United States.[86] That the high-yield bond markets of northern and southern Europe came into the US bond market's orbit in 1994 tended to surprise economists in these markets.

2. *Foreign (dis)investments and volatility*

The transmission of volatility in spillover studies is generally interpreted as the result of imitative trading behaviour across markets. Such contagion mechanism does not require the "contact" represented by interna-

[83]We also repeated the exercise including proximate market movements in the domestic volatility process. The broad pattern of the results is similar but fewer linkages can be detected. This suggests that the effects uncovered in the text operate in part through induced changes in the yield in domestic markets. See Annex VI for details.

[84]See Hamao, Masulis and Ng (1990).

[85]See also Koutmos and Booth (1995), who find that spillovers are stronger when the "news" in the original market is bad i.e. essentially, when the market declines. Given the link between downmarkets and volatility, this evidence is consistent with the message that volatility transmission is stronger when volatility is high.

[86]This would be the usual interpretation of the closer links between returns of US and UK shares, although their correlation is boosted by strong direct investment links that have no parallel in the case of bond markets.

tional flow of funds.[87] By contrast, we find evidence that such flows influenced volatility in several European markets during the bond market turbulence of 1994.[88]

The evidence

The procedure to test for the role of transactions by non-residents was to add their net investments in government bonds to the process for domestic implied bond yield volatility, possibly defined to include proximate market movements.[89] In keeping with the previous analysis, we allowed net sales and net purchases to have separate effects. Data were only available at a monthly frequency. We examined six countries: Germany, France, Italy, Canada, Belgium and Spain. Unfortunately, UK figures were only quarterly, too few data points for a meaningful estimation.

The evidence indicates that foreign activity was linked to increased volatility in all the G-7 European markets in our sample.[90] Moreover, the impact was asymmetric: withdrawals of funds had a considerably larger effect than their arrival, whose influence in some cases cannot be detected. The association between foreign sales and volatility is quite striking to the eye, at least in the case of Germany and France (Graph 23). For instance, non-residents liquidated over DM 13 billion of their holdings of German public debt securities in March 1994, a month in which implied volatility jumped by 4 percentage points.

The estimates suggest a sizeable impact (Table 10). Even allowing for the correlation between withdrawals and changes in domestic yields,[91] the

[87] At least one study has shown that the contagion around the 1987 stock market crash did not require such contact (Aderhold, Cummings and Harwood (1988)).

[88] Just as spillovers can occur without reinforcement from capital flows, so, too, the withdrawal of funds can destabilise a market without any spillover at work. At least on some accounts the Mexican fixed income markets were destablished in *late* 1994 by the withdrawal of foreign funds even as the immediate background conditions of the dollar bond market remained quite favourable. The Brady bond market did show high volatility in *early* 1994, consistent with a spillover, but experienced its worst gyrations in early 1995.

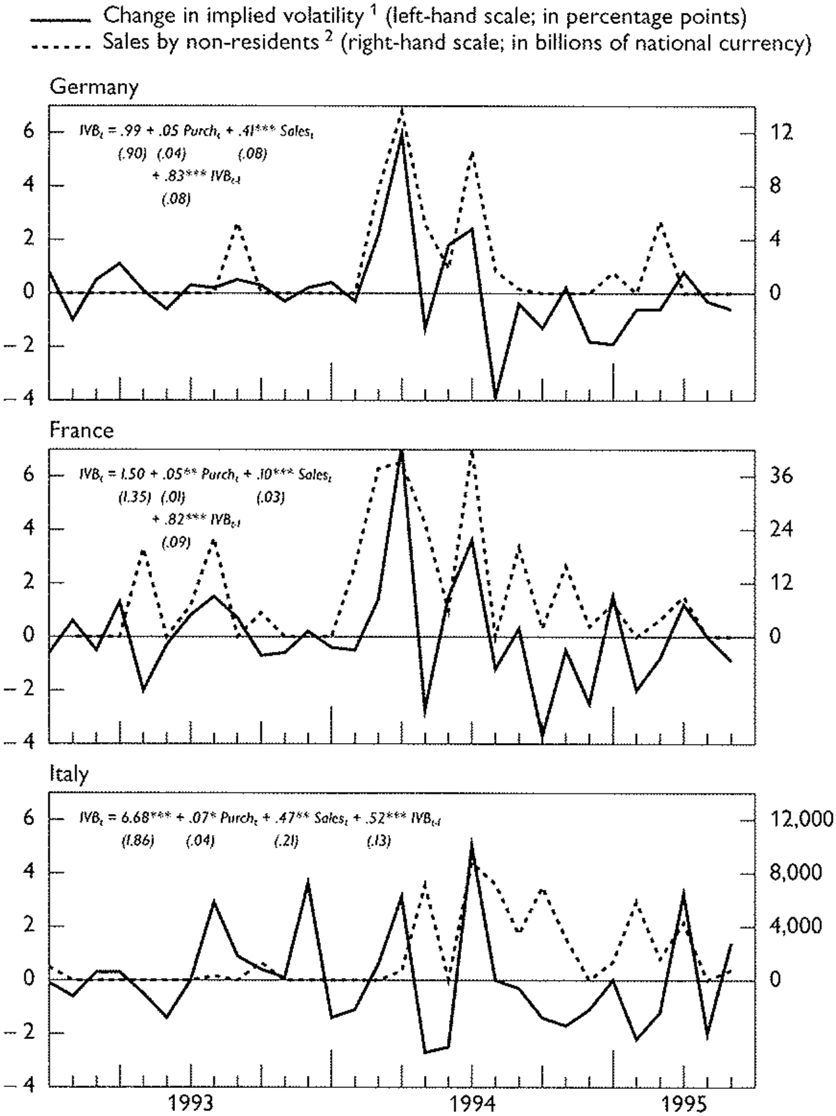
[89] See Annex VII for details.

[90] See Domanski and Neuhaus (1996) for evidence for Germany, and Stokman and Vlaar (1995) for evidence for the Netherlands.

[91] The fact that the link with foreign disinvestments survives the inclusion of increases in bond yields in the domestic market casts doubts on assertions that the link operates through the price effect associated with such sell-offs: bond prices fall and their relationship with higher volatility kicks in. (For an argument along these lines for Swedish shares, see Sellin (1994).) Only in the case of France the influence of foreign disinvestments cannot be clearly distinguished in statistical terms from that of market declines. The corresponding estimate of the impact, therefore, is subject to greater imprecision.

Graph 23

Bond yield volatility and bond sales by non-residents



[1] As defined in Graph 2. [2] Net sales are truncated at zero. For Germany, public sector DM-denominated bonds; for France, OATs and BTNs; for Italy, BTPs.

Sources: J. P. Morgan and central banks.

Table 10
Implied bond volatility and foreign disinvestment in selected bond markets, 1994

	Cumulative withdrawal[1] in billions of US$[3]	Change in volatility (in percentage points)			
		immediate impact per 1 US$ billion	cumulative[2]		memo item: 1994 vs. 1993
			to peak	in 1994	
Germany .	24.0	0.54	7.3	3.9	6.4
France . . .	27.0	0.44	7.4	3.8	5.8
Italy.	16.7	0.60	4.6	1.5	3.5

[1] February–June for Germany and France; March–July for Italy. [2] Effect of five-month period of sales. Based on AR(1) regressions for implied bond yield volatility (including, where appropriate, the impact of proximate market movements) to which foreign net purchases and sales are added separately. [3] Converted using the average exchange rate for 1994.

Sources: J.P. Morgan, national data and own estimates.

parameters indicate that the foreign liquidation of bonds equivalent to Fr.fr. 150 billion in France, DM 39 billion in Germany[92] and Lit. 27 trillion in Italy in the first half of 1994[93] had a peak impact on implied volatility in the corresponding markets of between 7 and 4 percentage points. Taken at face value, this bout of selling was associated with about two-thirds of the 1994 rise in average volatility in the French and German bond markets.

Foreigners appear to have disinvested from European bonds on derivatives exchanges as well as in the cash market. Foreign accounts are well-represented among holders of long positions on the European exchanges. For instance, MATIF officials estimate that one-half of long open positions are held by non-residents. The peaking in early 1994 and subsequent decline in open positions in the first half of 1994 (Graph 24), therefore, is consistent with the pattern of foreign activity observed in this cash market. It is a commonplace that cash transactions are less informative in a world of unseen derivatives positions. These data from the exchanges, at least, suggest that foreign disinvestment of cash holdings of

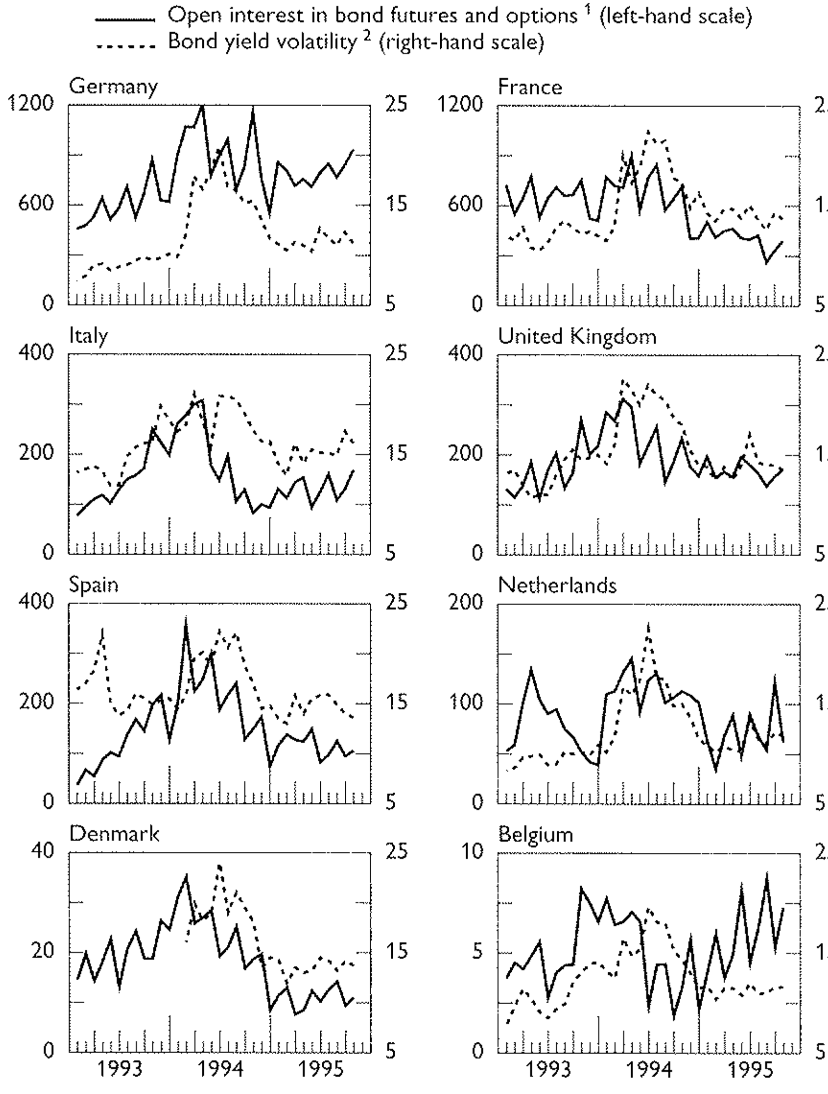
[92] Bundesbank analysts reckon that something like 5% of the German public bonds are held by residents through bank accounts in Luxembourg (see Bundesbank (1995)). But there was no fiscal reason for these fiscally motivated offshore resident holdings to be liquidated in early 1994.

[93] February to June for France and Germany; March to July for Italy.

Graph 24

Bond futures and option positions and bond yield volatility in Europe

At month-end



[1] In thousands of contracts. [2] In percentages, as defined in Graph 2.

Sources: Futures Industry Association, J. P. Morgan and own calculations.

European bonds in 1994 actually *understate* the change in foreign positions, on and off-balance-sheet.

Interpreting foreignness

But why should the identity of the seller matter? In our view, the link between crossborder flows and bond volatility does not derive from foreignness per se. In particular, it would be a mistake to say that exchange rate risk distinguishes foreign from domestic investors. Nowadays non-residents not only *can* hedge the currency risk out of their bond portfolios, they also do so with some regularity. The mechanics are various, ranging from forward sales of the currency in which the bond is denominated to financing the long bond position in the repo market, so that an overnight liability in the relevant currency matches the bond.

Some evidence can be put forward to establish the regularity of hedged foreign positioning in European bonds in 1993. Interviews in New York in the autumn of that year suggested that major international bond managers were using hedged positions as the base investment, with perhaps a separately considered or even separately managed currency overlay. It is striking that no less than $182 billion of short-term funds flowed through interbank channels out of European money markets in 1993, when foreign investment in European government bonds reached a similar figure (Table 2, first memorandum item).

If exchange rate risk need not be the distinguishing criterion, the more general difference remains that foreign investors in bond markets nowadays are typically leveraged. The combination of a long bond holding and a short overnight cash position in the same currency, a yield-curve play, is a synthetic asset. In the hands of a non-resident, it matches no liability, in contrast to an outright bond holding by a domestic insurance company or pension fund with long-duration liabilities in domestic currency. Partly in consequence, foreign investors are more prone to seek to maximise total returns in the context of mark-to-market accounting than, say, a domestic insurance company.[94] Much of the leveraged purchases of bonds identified in Section II were undertaken by securities firms operating in London (Table 2).[95]

[94]For some evidence, see e.g. Frijus, Kleinen and Quix (1995).

[95]Not included in these figures would be hedge funds resident in the UK (or in the Caribbean).

Such leveraged investments have been facilitated in recent years by specific developments in the credit markets. In particular, the establishment and rapid expansion of repurchase markets on the continent in the 1990s have permitted foreign investors to borrow beyond the limits on their access to *unsecured* credit for the financing of their bond portfolios.[96] In other words, by selling temporarily their bond holdings and using the proceeds to finance further purchases, firms could swell their portfolios on the back of short-term financing well beyond previous limits (Table 2, second and third memorandum items).[97] The development of options markets has had a similar effect.

Interpreting the asymmetry of foreign investment

The heavy reliance on leverage by non-residents can partly help to interpret the impact of foreign sales and the asymmetry with respect to purchases. Leveraged investors constrained by mark-to-market accounting would find it difficult not to sell as prices decline steeply. In the process, their rapid withdrawal would put liquidity under pressure. The asymmetry of the effect could then be explained by the different speed of purchases and sales, reinforced by the asymmetry in issuance and retirement policies already discussed in Section II.[98] Foreign investors reportedly build up their positions gradually, but in stress situations sell at a more rapid pace. On the way in, they help to absorb the growing supply of government debt: if they wish to stake out a large position, they can simply wait until the next auction and bid a basis point or two above the market. On the way out, however, if they wish to sell an equivalent large stake, the government is not there to take the other side of the trade. They have to find a local bank or insurance company willing to increase its exposure in an uncertain market. The contrast between these two situations clearly illustrates how the impact on liquidity need not be symmetric.

[96]See Bank for International Settlements (1995b).

[97]The proposition that foreign investors are more leveraged is not a timeless or placeless assertion, but *a view of the bond market in the 1990s*. In the mid-1980s, by contrast, one of the largest net flows in the world bond market was the purchase of US Treasuries by Japanese life insurers. These investments came to be hedged to a varying extent, but they were not leveraged in the sense of being financed with short-term funds on the basis of a fairly slim net worth.

[98]In our interviews, market participants pointed to this as a possible explanation.

Conclusions

Let us summarise some of our key findings. Bond yield volatility typically rises in response to downward movements in the bond market but over time tends to revert to its mean. The long-term level of volatility responds to the success of price stabilisation policies and reflects difficulties in fiscal management. Variations in bond market volatility are associated with variations in money market volatility. There is, however, little evidence that uncertainty about fundamentals such as inflation, growth, fiscal balances or the short-term conduct of monetary policy lay behind the turbulence in bond markets in 1994. During that episode, bond volatility communicated itself across national boundaries in a hierarchy of markets. And the withdrawal of foreign investments appears to have raised volatility in a number of large European markets. What are the lessons of these findings for policy?

1. Lower volatility from lower inflation

Monetary policy plays an important role in setting the base level of bond market volatility. Price performance that damps out inflationary expectations and permits bond yields to fall will have a welcome side-effect on volatility. The clear conjunction of high expected inflation, record high interest rates and record high volatility in the US bond market fifteen years ago highlights the importance of inflation. In the United Kingdom, too, volatility peaked in the period of high and volatile inflation. And the coexistence of low volatility in the core European markets alongside the higher volatility in their high-yielding counterparts points in the same direction. Wringing the inflation out of an economy can calm the bond market.

It is worth emphasising the factors that did *not* accompany the highest volatility in the history of the US market. This volatility did not require well developed markets for bond futures and options, new forms of leveraged investments and massive variations in holdings by foreign investors. Bond futures were in their infancy and option strategies had progressed little past stop loss orders. Similarly, international transactions in bonds were negligible by today's standards and dollar bonds outstanding and trading bulked much larger in world outstandings and transactions than they do now. It is the policy errors that allowed inflation expectations and nominal rates to rise to very high levels that must bear responsibility.

2. *Fiscal policy*

Fiscal policy, too, can affect bond yield volatility. To the extent that reducing fiscal imbalances can help to bring down the level of yields, it will also have a moderating influence on their volatility. Cross-sectional evidence is broadly consistent with this link.

In the time series, the link between changing apprehensions about fiscal policy and bond yield volatility is only visible in the country of our sample with the most serious fiscal imbalance, Italy. Changes in the swap spread, our high-frequency measure of concerns about default risk, help to explain movements in bond volatility. The use of the spread as a meaningful indicator of fiscal concerns is limited to high-debt Italy, however.[99] Thus, our findings are limited by our measure, not necessarily by the import of fiscal policy.

3. *Limiting money market volatility*

Several pieces of evidence suggest that bond volatility reflects money market volatility: the coincidence of record-high money and bond volatility in the period of non-borrowed reserves targeting in the United States; regression analysis of realised money market volatility and implied bond volatility for thirteen industrialised countries; and the relationship between US dollar LIBOR cap implied volatility and US implied bond yield volatility. The pass-through, however, looks to be modest, averaging some 5% within one week. Thus, for most of the markets examined, the variation over time in money market volatility explains a fairly modest portion of the short-term movements in bond volatility. On the other hand, its wide range of variation across countries accords it a larger role in explaining international differences. The contrast between the placid German money market and the more volatile Japanese one can account

[99]The swap spread is normally a compound of the default risk of the private sector and supply and demand for fixed rate funds by private liability and asset managers. Only at extreme levels of public indebtedness can the level and change of the swap spread be interpreted primarily as a reflection of government default risk. Indeed, Italy is the only country where the swap spread is typically positive, i.e. where the government has to pay a higher rate on its borrowing than prime quality private corporations (see Giovannini and Piga (1992)).

for a significant portion of the gap in bond yield volatility between the two countries.

The role that central banks play in the determination of money market volatility remains a topic for research. We documented how the Federal Reserve's 1 percentage point reduction in official interest rates in December 1991, an unusually big move, was associated with a sharp rise in the implied volatility priced into 3-year LIBOR caps. But we did not examine more generally the importance of central bank operating procedures and strategies for money market volatility.[100]

4. *Limited power of fundamentals*

While fundamental economic factors help to set the background level of volatility,[101] their impact on its short-term movements is hardly apparent, at least when judged on the basis of experience since 1992. We did not find much evidence that uncertainty about fundamentals such as inflation, growth, fiscal policy or the short-term conduct of monetary policy could explain the major rise in volatility during 1994.[102] Survey measures of the volatility and dispersion of market participants' inflation and growth forecasts do very poorly. With the notable exception of the United States, 1994 subsidence of money market volatility, a possible index of uncertainty about monetary policy, should generally have been expected to calm bond markets. And only in Italy were there signs of an influence of fiscal policy.

Fundamentals could indirectly explain the turbulence in bond markets in 1994 through their impact on the *level* of yields: we found fairly generalised evidence that bear markets tend to raise volatility considerably. Yet independent analysis suggests that it is difficult fully to account for the

[100] Other questions of a more specific nature are raised by the cross-sectional and time series variation in money market volatility. To what extent should small open economies that seek to limit the range of fluctuation of their exchange rates expect to experience periodic bouts of money market volatility connected to exchange rate strains and what are the implications for bond volatility? Is part of the volatility of bond markets at turning points in monetary policy related to the volatility of money markets at such turning points? If one reason for the high volatility of the Japanese bond market is the relatively high volatility of its money market, why is the Japanese money market so volatile in the first place? Can one interpret the Bundesbank's switch from flexible to fixed rate tender repos in July of 1994 as a response to unwanted money market volatility? Was the switch effective in calming the market? See Timmermans, Delhaz and Bouchet (1995).

[101] See Morton (1996) for cross-sectional evidence.

[102] See Borio and McCauley (1995).

sharp rise in bond yields in 1994 in terms of fundamentals either. If so, good policies are a necessary, but not a sufficient, condition for low volatility.

5. *International spillovers, capital flows and the power of markets*

International linkages played a major role in the 1994 bond market turbulence. Spillovers of volatility grew in size and geographical reach, with volatility emanating initially from the United States. And our evidence indicates that rapid foreign disinvestment rocked at least three large European bond markets, viz. Germany, France and Italy. Thus, the bond volatility of 1994 seems to have reflected less investors' unjustified fears of inflation[103] than their justified apprehensions about other investors' activity.

In general, outflows appear to have had a larger impact than inflows. In our view, this relationship indicates that leveraged players constrained by mark-to-market accounting are nowadays over-represented among foreign investors in bond markets. As prices decline steeply, such investors quickly come under great pressure to sell at a time of unreceptive market conditions, straining market liquidity.

International capital flows have recently been larger and more volatile than at any time in living memory. If one accepts the proposition that the comings, and especially the goings, of international investment can drive up volatility even in the relatively broad and deep European government bond markets, what is the appropriate policy response?

One perspective is that good policies will tend to yield lower bond volatility over time, but that a market open to foreign investment is vulnerable to an additional source of volatility in global downturns. Good policy should perhaps be seen as producing greater stability on average over time, but not as offering any insulation from ambient international turbulence. In particular, developing countries that in the past have maintained polices to bar or to slow the influx of foreign fixed-income investors need to recognise that while opening their money and bond markets promises clear benefits, there may be associated costs too. Care and thought should guide such policies (BIS (1995a)).

[103] See Goodfriend (1995).

One can also interpret recently announced changes to the gilt market in the light of our findings. Until the end of 1995 UK officials restricted the access of investors, including international investors with a leveraged *modus operandi*, to the repurchase market. In January 1996 the authorities finally opened the gilt repo market.[104] This opening was explicitly premised on the view that the repo financing mechanism could bring in new investors, who had the potential of reducing the government's borrowing costs. Given the swings associated with repo-financed investors, especially foreign ones, the UK policy can be read as accepting a possibly *more volatile* demand, and its potential for an asymmetric knock-on to volatility, in exchange for a *higher average* demand, with its savings in debt servicing costs.

6. *Debt management: sales of embedded options and the timing of issues*

Our findings confirm the predictable, mean-reverting nature of bond volatility. This feature helps to highlight from a new perspective two issues in debt management. The first is the fairly new one of government sales of options, the second is the very old question of the timing of issues.

Government sales of volatility?

Given the predictable nature of volatility, when debt managers observe market participants placing a high value on it, why should they not sell the right to buy a bond at a high price (or to sell one at a low price)? Debt managers are in fact already doing so, albeit in the less transparent form of options embedded in bonds.[105, 106]

In 1994, for instance, UK debt managers recognised a period of high bond market volatility, and sold an option at a correspondingly high price in the form of a convertible gilt.[107] Since volatility was near its 1994 peak

[104] See UK Treasury and Bank of England (1995).

[105] The report on the macroeconomic implications of derivatives by the central banks from Group-of-Ten countries (the "Hannoun Report") considers the use of derivatives by central banks very cautiously (see BIS (1994a), pp.43-52). As advisors to their treasuries, however, some central banks have not altogether refrained from in effect selling options.

[106] For years, the US Treasury made its 30-year bonds callable after 25 years; several years ago it exercised one such option and then ceased to attach the calls – perhaps after dealers had learned of the need to price them in.

when the issue was announced in mid-May, the market placed an unusually high value on the option feature, which found expression in a reduction in the required yield.[108]

Volatility and the timing of debt issues

The mean-reverting property of volatility also bears on the long-standing question of regularity *versus* opportunism in debt management. In one camp is the US Treasury, which was persuaded in the 1970s and has gone on to persuade a number of foreign governments that a pre-announced schedule of bond auctions is the best long-term approach to selling paper.[109] The US Treasury's rationale emphasised three points: the unwelcome effect of unexpectedly large debt offerings on the market, the benefit of allowing market-makers and investors to plan better, and the practical difficulty of successfully timing the market. In another camp is, say, the World Bank, which as issuer of paper takes its opportunities as it sees them.

Whatever one thinks of the ability to time issues in terms of choosing periods when rates are comparatively low, a treasury has a well-founded

[107] The Bank of England acting on behalf of the UK Treasury sold a 3-year note yielding 7% but convertible on four dates into a 17-year bond yielding 9% (Bank of England (1995), p. 24). Such a bond can be considered a compound of a medium term plain vanilla bond and four warrants, or long-dated options, to buy a long bond. The Bank of England (1995), p. 5, reports laconically: "A convertible stock – short into a long – was auctioned in the highly turbulent conditions of May 1994 when uncertainty was high; implied volatility on the option on the gilt future averaged 12.4% in May 1994, and was 11.5% on the day of the auction, compared with an average of 7.8% in 1993-4."

[108] Notwithstanding this apparent success in selling high, the recent joint Treasury/Bank of England debt management review stated without argumentation that "Convertible gilts will not form part of the regular issuance programme, but may be sold in exceptional circumstances". See UK Treasury and Bank of England (1995), p.12.

[109] For an account of the US Treasury's rationale for "regularization," see Baker (1979). The chief economist at Salomon Brothers agreed with what he called the monetarist view; see Kaufman (1973). "The present policy of engaging in large operations at discontinuous intervals in effect forces the government to speculate on the course of market rates"; "the present policy of issuing a wide variety of securities and seeking to tailor their terms and date of issue to the market... is a fertile source of confusion, uncertainty, and instability." (Friedman (1960), pp. 63–64). The Bank of England has tended to approach the US Treasury's *modus operandi* over time. Thus the Bank of England (1993), p. 22, stated its goal as "to ensure reasonable predictability in the issue of stock, which helps to avoid large and arbitrary price movements, while retaining flexibility to respond to market conditions." Her Majesty's Treasury and the Bank of England (1995), pp. 13–16, more recently turned further away from such flexibility, that is "uncertainty" and "unpredictability," in favour of pre-announced auctions. Tap sales are to fall from around 57% of conventional government bond issuance in 1992–93 to about 19% in 1994–95 and to not more than 10% in normal circumstances in the future.

hope of avoiding highly volatile markets: the well-documented tendency for volatility to revert to its mean grounds this hope. Since it is reasonable to assume that a government holding an auction in a particularly volatile market pays for the privilege,[110] the commitment to a pre-announced fixed schedule of bond issuance risks doing away with valuable flexibility. For instance, the action of certain European governments in delaying debt issues in the middle of 1994 might be justified as a response to high volatility that was predictably going to fall. The possible savings from identifying ex ante less volatile market conditions into which to issue are more readily measured[111] than the as yet unquantified benefits of regularity of issuance.

[110] A bid for government bonds resembles a put option written by the primary dealer to the Treasury – short-lived to be sure, but nevertheless obligating the dealer to buy at a schedule of prices without providing any assurance of purchase. See Smith (1976).

[111] The cost of the implicit put depends, inter alia, on the prevailing short-term volatility and on the lag between the bid and the announcement of the results, at which time an option-type position turns into a simple cash position. To obtain some idea of the cost of holding an auction in a particularly volatile market, consider the cost of one-day at-the-money options on Treasury bonds. The premium ranges around 8/32s normally, but on the day before the employment report reaches 12/32s (see Ray (1993), p. 199). The difference, 4/32s or 0.125%, provides an upper limit to how much a dealer might shave his bid for Treasury bonds to compensate for a more volatile market environment.

Annex I – The data and their advantages

The basic database consists of weekly observations on implied yield volatilities for three-month over-the-counter (OTC) at-the-money options on 10-year benchmark government bonds in thirteen major markets as quoted by a leading market maker, J.P. Morgan. This annex considers two key issues: the choice between *yield* and *price* volatility and the relative merits of OTC and exchange-traded implied volatilities.

Yield vs. price volatility

The two most widely used measures of volatility are *price* and *yield* volatility, respectively the (annualised) standard deviations of daily[112] percentage changes in the price and in the yield to maturity of the bond. The measures would be equivalent if the cash-flow stream associated with the security was a unique payment at maturity ("zero coupon bond"). Otherwise, the payment of intermediate coupons drives a wedge between them.

Which measure is more appropriate depends on the specific question asked. Price volatility is the relevant concept when assessing the potential gain or loss from holding a bond. Thus, "value-at-risk" models for the evaluation of the market risk associated with investors' portfolios rely on estimates of price volatility.

Nevertheless, for comparisons of volatility across national markets and time the measure based on the yield is more useful. Conceptually, it is the volatility of individual discount factors for the coupon payments that is taken as the primitive notion affecting yield and price volatility. And in contrast to price volatility, the link between yield volatility and the volatility in those discount factors is not affected by coupon payments. For example, if short-term rates are assumed constant over time, the yield to maturity and the discount factor coincide, regardless of the coupon payment. Indeed, in "value-at-risk" models price volatilities are

[112] Note that, when percentage changes in the yields (prices) are correlated over time, the length of the interval over which the individual changes are measured will affect the (annualised) measure of volatility. For an elaboration of the implications of this result, see Cohen (1995).

generally *derived* from yield volatility: it is the yield volatility process which is modelled and taken as given.[113]

In what follows, therefore, we will focus on yield volatility. As a concrete example, if a bond yielding, say, 10% moves by 10 basis points, the yield edges by 1%. Annualising this change by the square root of 256, the standard number of working days in a year, gives a yield volatility of 16%.[114]

OTC vs. exchange-traded options

The pricing indications derived from OTC options have two overriding advantages over their exchange-traded counterparts. They cover a broader set of markets: they exist also for government bonds that are not exchange traded. And they are quoted for the *same maturity at every observation*. This is not possible for implied volatilities on exchange-traded contracts, as these contracts exist only for fixed calendar dates. Successive quotes on the same contract thus differ whenever implied volatility varies across maturities. In other words, unless volatility is anticipated to remain constant, the mere passage of time introduces an extraneous source of variation.[115] While interpolating techniques have been developed to deal with this problem,[116] the constant-maturity aspect of the OTC quotes avoids them altogether.

[113] See J.P. Morgan (1994) p. 81. Price volatility is equal to yield volatility times yield times modified duration, where the latter is a weighted average of the bond's cash flows with weights being a function of time (see e.g. same source). As an illustration of the difference between yield and price volatility, consider the comparison between the benchmark US Treasury bond and its Swedish counterpart in mid-September 1995. The US security had a coupon of 6.5%, the Swedish instrument one of 6.0%. Since krona yields exceeded dollar yields by a sizable margin, the Swedish bond sold at a heavy discount; the US security, by contrast, traded close to par. As a result of the deep discount, the Swedish bond approached the long duration of a zero coupon bond. Measured in terms of yield, the implied volatility of the US security was higher, 18.2% against 16.5%. In terms of price volatility, however, the Swedish bond appeared to be considerably more volatile, 10.3% against 8.2%. Yield volatility abstracts from accidental differences in discount or premium pricing of benchmark bonds.

[114] Having taken the yield as the reference variable, a further possible choice is that between volatility measured in *percentage* changes and in *basis point* changes. Theory can hardly guide this choice. Option models, for instance, assume both kinds of interest rate processes. Our choice of percentage changes is largely driven by the original data used, which were expressed in this way (see below). In addition, most extant academic research uses this measure (e.g. Singleton (1994) and Figlewski (1994)).

[115] As an illustration, consider the following typical example. Since volatility is mean-reverting, a period of comparatively high volatility is typically one with a downward-sloping term structure of volatility. In such an environment, implied volatility in a given contract rises as the contract approaches expiration quite apart from any movement in the term structure.

[116] See e.g. Canina and Figlewski (1993) for the adjustment as applied to exchange-traded equity options.

Table AI.1

Turnover of government bond options, April 1995

In billions of US dollars per day

	OTC[1]	Exchange-traded
US Treasuries	1.7	11.8
Japanese Government Bonds	6.0	11.1
German Bunds	0.7	4.5
Other	1.8	6.1[2]
Total	10.2	33.5

[1] Over-the-counter single-currency interest rate options on traded securities. [2] French, Italian, British, Spanish, Dutch, Belgian and Australian government bonds.

Sources: BIS (1996a), Chicago Board of Trade, Tokyo Stock Exchange, Singapore International Monetary Exchange, London International Financial Futures and Options Exchange, Deutsche Terminbörse, Marché à Terme International de France, Mercato Italiano dei Futures, Meff Renta Fija and Sydney Futures Exchange.

A potential disadvantage of the data is the use of dealers' quotations, and from a single market-maker, rather than of the more transparent transaction prices collected by the more active exchanges (Table AI.1). This potential disadvantage, however, should be kept in perspective.

At the outset, recall that financial markets have repeatedly confronted the problem of the reliability of OTC quotations in the past. The most famous example is the London Interbank Offered Rate (LIBOR) for bank deposits, which, just as an OTC option contract, can expose the buyer to the selling bank's credit risk. Big syndicated loan contracts with interest rates tied to LIBOR will typically specify the five leading banks whose quotations are to be averaged.[117] The difference between an unquestioned acceptance of LIBOR and of our OTC quotations thus reduces from the principle of using OTC prices to the practical question of whether one can rely on *one* dealer's prices.

Those in charge of monitoring the accuracy of the portfolio valuation of an institution's own dealers typically use those of competitors as a

[117] As an additional illustration, a glance at this morning's Wall Street *Journal* will show prices for US Treasury bonds, notes and bills, with quotations supplied by the Federal Reserve Bank of New York for ordinary paper and by Bear Stearns for strips. The reason is that the US Treasury's facilitation of strips was not accompanied by the New York Fed's extending its price collection. The New York Fed obtains the quotations by calling a number of primary dealers.

benchmark.[118] It is therefore natural to do the same in our case. A comparison of the J.P. Morgan quotations with scattered ones from Midland Montague, Hong Kong Banking Corporation's London affiliate, was reassuring. Given differences in the timing of the quotations and the need to convert price into yield volatility through a standard approximation, the remaining small discrepancies indicated that the J.P. Morgan quotations were a satisfactory basis for the analysis.[119]

[118] For instance, an option trader was convicted of fraud for entering flattering volatilities into the programme that evaluated his option book. How were those who monitored his performance to check his inputted volatilities? The procedure was to call Bankers' Trust. See Millman (1995).

[119] For the French OATs, we compared volatilities on nine days in the first four months of 1995 and found that the difference between the two sources averaged 0.67 basis points and ranged between 0.06 and 1.58. The difference amounted to 4.6% of the level of volatility. Given that there is little consistency regarding the time of day at which the Midland Montague volatilities are reported, while the Morgan data are Thursday closes, these results are not too disquieting.

Annex II – Persistence in implied bond yield volatility

This annex considers in more detail the statistical evidence on the time-series properties of implied bond yield volatility. It focuses mainly on the autoregressive properties of this variable as the basis for subsequent tests aimed at evaluating its relationship with other economic factors. Two conclusions are reached. The balance of the evidence indicates that bond yield volatility is best regarded as a stationary series. In addition, one autoregressive lag for all countries appears to be a satisfactory basis for subsequent tests.

A key question about the time-series properties of an economic variable is whether it is stationary. Stationarity implies that over time the variable tends to fluctuate around its mean value, possibly a time-varying one. In the context of the present paper this issue is relevant for two reasons. First, it tells us something about our ability to predict the variable over different horizons. If a variable is stationary then a good guess of its future value is its long-term average. In particular, non-stationarity actually implies that the expected forecast error based on the past history of the variable tends to infinity as the horizon is lengthened indefinitely. Second, the issue is relevant when assessing the relationship between variables. If two non-stationary series are taken in consideration, certain statistical tests are generally biased towards finding a significant relationship even when none exists. A proper analysis would then have to examine whether the variables in question are co-stationary, i.e. roughly speaking, whether a weighted average (linear combination) of the variables is a stationary series.

When performed on the individual time series, Augmented Dickey Fuller tests (ADF) were often unable to reject the hypothesis of non-stationarity (i.e. of the presence of a stochastic unit root). Only in the cases of Italy and Spain was there strong evidence of stationarity; the evidence was weak for the United States and France. This was so regardless of the number of own lags included in the tests (one to four) and of whether a deterministic linear trend was assumed or not (see Table AII.1 for a sample of the results). The first difference in implied volatility, by contrast, easily passed the test (not reported). Taken at face value, this evidence would suggest that implied volatility is generally non-stationary and its first difference stationary.

Nevertheless, there are grounds for doubting that implied volatility is actually non-stationary. As is well known, the power of unit-root tests can be quite low. Given that the property under scrutiny relates to the very long run behaviour of the series, two years of data may not provide a sufficiently large sample to reach conclusions with much confidence. In particular, it may be difficult to distinguish a non-stationary series from one that, while stationary, reverts relatively slowly to its mean. Non-stationarity not only runs counter to much of the econometric literature on the treatment of asset price volatility i.e. various variants of GARCH estimation techniques. It is also a property not supported by the analysis of the forecasting performance of bond yield volatility estimates over different horizons (Figlewski (1994)) and disbelieved by market participants.

Given the limited length of the sample period, one way to increase the statistical power of the tests is to consider the cross-sectional information. This can be done in at least two ways. The first is to estimate the relationship between implied volatility and its own past *jointly* across countries, taking into account the information contained in the cross-country correlation of residuals (through a SURE technique). The second is to go one step further and to estimate a *single* weighed average autoregressive coefficient by pooling the data.

Table AII.1 reports the SURE estimates of an autoregressive process with one lag (AR(1)) based on weekly data. The estimates provide some support for the view that implied bond yield volatility is stationary. The point-estimates of the autoregressive coefficient are somewhat lower than the single-equation ones, typically by between 5 to 10 percentage points. The corresponding values now range from 0.71 to 0.92, against a value of 1.0 associated with non-stationarity. The precision of the estimates, as indicated by the standard errors, changes little. As a result, *if* one simply applied the traditional ADF marginal significance values to the above set of numbers, non-stationarity is rejected, generally at the 1% level. Admittedly, this experiment should be taken as illustrative only: we do not know how the ADF statistic should be adjusted to cope with the SURE estimation technique. Nevertheless, the results appear to shift the balance of the evidence in favour of stationarity.

Pooled estimation strengthens this shift. When the autoregressive parameter is constrained to be common to all countries, the estimate ranges from 0.89 (SURE) to 0.92 (OLS) with a standard error of around

Table AII.1

Comparing single and joint-equation estimation

Weekly data

	Single estimation			Joint estimation			Sample begins on[1]
	Coeff.	SE	DF t-stat.[2]	Coeff.	SE	DF t-stat.[3]	
US	0.90	0.04	–2.69*	0.85	0.03	–4.54***	31.08.92
JP	0.94	0.03	–2.07	0.92	0.03	–2.90**	31.08.92
DE	0.96	0.02	–1.63	0.89	0.02	–6.16***	31.08.92
FR	0.90	0.04	–2.71*	0.81	0.03	–6.41***	31.08.92
UK	0.96	0.02	–1.72	0.91	0.02	–4.42***	31.08.92
IT	0.84	0.04	–3.65***	0.78	0.04	–5.95***	31.08.92
CA	0.95	0.03	–1.94	0.91	0.02	–3.87***	31.08.92
BE	0.94	0.03	–2.11	0.90	0.02	–4.43***	31.08.92
NL	0.97	0.02	–1.63	0.91	0.02	–4.97***	31.08.92
ES	0.76	0.06	–4.20***	0.72	0.05	–5.57***	16.11.92
DK	0.92	0.05	–1.62	0.84	0.04	–4.01***	14.02.94
SE	0.94	0.04	–1.44	0.90	0.04	–2.87*	14.02.94
AU	0.88	0.06	–2.00	0.83	0.06	–3.07**	21.03.94

	Coeff.	SE	DF t-stat.[3]	χ^2	Prob. (%)
Pooled: All					
OLS	0.92	0.01	–8.16	22.0[4]	3.8**
SURE	0.89	0.01	–9.38	37.9[4]	0.02***
Pooled: All, excl. Italy and Spain					
OLS	0.94	0.01	–6.49	6.62[5]	76.0
SURE	0.95	0.01	–7.81	14.6[5]	15.0

[1] The sample period ends on 22.05.95 for all countries. [2] Relates to the test without additional lags and no time trend. The results were similar in the other cases. [3] t-statistic on lagged level of volatility when the regressor is written as a first difference. The significance level is only indicative since it applies the same thresholds as in the single-equation case. [4] 12 degrees of freedom. [5] 10 degrees of freedom.

0.01 in both cases. Subject to the qualifications just raised about the required adjustment to the ADF test, the corresponding t-statistics are even higher than those resulting from the unrestricted SURE estimates. True, the restriction that the autoregressive parameter is the same for all countries is rejected. But this reflects only the inclusion in the sample of the two countries with the *lowest* autoregressive coefficient, viz. Italy and Spain, for which stationarity was already accepted on the basis of the single equation estimates. In fact, if these two countries are excluded, the

Table AII.2

Diagnostic statistics for autoregressive equations[1]

Weekly data

	$\bar{R}^2$	SER	LM(4)[2]	Q(12)[2]	ARCH(4)[2]	WHITE[2]
US	0.81	0.77	72.7	73.3	96.5	67.8
JP	0.87	1.66	39.5	59.0	2.8**	0.0***
DE	0.93	0.95	0.1***	0.0***	9.7*	9.6*
AR(3)[3]	0.93	0.91	8.1*	31.1	82.4	53.5
FR	0.81	1.51	0.2***	0.0***	1.3**	0.1***
AR(2)[3]	0.84	1.43	78.8	9.6*	23.4	8.4*
UK	0.92	0.93	27.2	22.6	0.3***	3.1**
IT	0.73	1.42	94.7	64.6	49.5	90.3
CA	0.90	0.89	33.9	13.7	25.7	6.3*
BE	0.90	0.98	19.6	22.0	0.0***	0.0***
NL	0.94	0.81	53.2	78.5	0.0***	6.5*
ES	0.57	1.94	1.1***	29.1	0.0***	0.4***
AR(2)[3]	0.61	1.86	89.5	48.3	40.2	59.5
DK	0.84	1.47	17.8	46.2	6.0*	14.6
SE	0.88	1.52	84.8	25.0	79.2	58.1
AU	0.77	1.52	82.3	10.2	88.7	1.6**

[1] AR(1) unless otherwise stated. [2] Marginal probability of significance. [3] The sum of the lagged coefficients for Germany, France and Spain is 0.98, 0.85 and 0.84 respectively.

restriction passes comfortably at the traditional levels of significance while the Dickey-Fuller t-statistic remains quite high.[120] On balance, therefore, the judgement that implied bond yield volatility is stationary appears to be justified. This is further supported by the sizable fall in the point estimates of the coefficient using month-end data (see below).

The main criterion used for the choice of autoregressive lag length was to select a parsimonious representation and to limit tailoring to individual countries so as to highlight the comparatibility of the results. Tables AII.2 and AII.3 summarise the main results of assuming a single lag for, respectively, weekly and monthly data.

Looking at the weekly data first, the estimates indicate that an AR(1) process yields a high explanatory power (judged from the $\bar{R}^2$ statistic). In addition, there is generally little sign of residual serial correlation, which

[120] Actually, dropping Spain is sufficient to have the restriction accepted on the basis of the OLS estimates.

Table AII.3

Results of autoregressive equations

Monthly data

	Coeff.	SE	$\bar{R}^2$	SER	LM (4)[1]	Q (12)[1]	ARCH (4)[1]	WHITE[1]	Sample begins in[2]
					Month-end				
US	0.72	0.12	0.50	1.28	38.8	24.7	40.6	42.4	Sept. 1992
JP	0.82	0.13	0.64	2.97	83.8	81.8	98.9	74.0	Sept. 1992
DE	0.87	0.10	0.79	1.63	29.0	65.8	90.8	44.3	Sept. 1992
FR	0.83	0.10	0.68	1.88	74.4	81.3	60.3	70.3	Sept. 1992
UK	0.83	0.07	0.68	1.89	71.0	87.3	94.7	45.2	Sept. 1992
IT	0.70	0.09	0.53	1.82	52.7	63.7	94.3	42.8	Sept. 1992
CA	0.80	0.11	0.64	1.69	4.6**	43.6	75.1	2.3**	Sept. 1992
BE	0.85	0.08	0.77	1.49	69.6	78.0	13.4	40.3	Sept. 1992
NL	0.85	0.13	0.76	1.74	36.2	60.8	10.7	4.2**	Sept. 1992
ES	0.59	0.16	0.33	2.24	73.1	84.3	98.8	4.9**	Dec. 1991
DK	0.73	0.17	0.45	2.69	8.4*	37.3	99.7	38.9	March 1994
SE	0.72	0.19	0.47	3.32	34.1	56.4	5.2*	58.6	March 1994
AU	0.49	0.18	0.23	3.14	56.1	98.5	46.5	43.7	April 1994
					Month-average				
US	0.81	0.10	0.65	0.97	37.5	43.4	17.6	89.7	Sept. 1992
JP	0.88	0.09	0.76	2.18	38.8	13.0	99.9	60.2	Sept. 1992
DE	0.91	0.07	0.86	1.32	38.7	48.6	92.0	61.2	Sept. 1992
FR	0.89	0.08	0.79	1.53	3.0**	31.4	25.3	65.3	Sept. 1992
UK	0.91	0.07	0.81	1.39	29.9	40.8	21.4	24.0	Sept. 1992
IT	0.62	0.10	0.42	1.93	59.6	79.8	51.6	47.8	Sept. 1992
CA	0.83	0.09	0.67	1.51	19.2	46.5	43.1	37.1	Sept. 1992
BE	0.88	0.09	0.82	1.31	49.6	45.0	75.1	8.4*	Sept. 1992
NL	0.91	0.08	0.96	1.27	16.2	13.3	77.5	31.6	Sept. 1992
ES	0.68	0.12	0.44	1.98	32.6	7.7*	55.3	61.3	Dec. 1991
DK	0.81	0.13	0.59	2.29	2.9**	13.0	65.2	21.9	March 1994
SE	0.81	0.21	0.61	2.74	48.3	74.7	79.6	1.4**	March 1994
AU	0.78	0.18	0.59	1.87	23.9	18.5	81.7	14.2	April 1994

[1] Marginal probability of significance. [2] The sample period ends in May 1995 for all countries.

would signal the presence of systematic unaccounted factors. The exceptions are Germany, France and Spain. Extra lags were significant only in these three cases.[121] Although they helped to eliminate the signs of serial

[121] And in Denmark, where, however, they introduced serial correlation.

correlation, they had otherwise little impact on either the sum of the coefficients or the explanatory power. The choice here was to report the results using one lag for three reasons: this procedure makes cross-country comparisons of parameters easier; the results proved invariant to the inclusion of further lags; and for France and Germany these lags actually dropped out following the introduction in the regressions of the terms capturing directionality, the key controlling factor in subsequent tests.

The picture is similar in the case of the monthly estimates (Table AII.3). Whether month-end or month average figures are used, there are only very few signs of serial correlation and little to be gained by including extra lags.

With regard to heteroskedasticity, by contrast, the estimates indicate that it would be imprudent to ignore its presence at least at the weekly frequency. For the sake of consistency, we report White heteroskedasticity-consistent standard errors throughout.

Annex III – Estimating directionality

This annex examines the relationship between implied bond yield volatility and bond market movements. It asks three questions. First, do movements in bond yields add explanatory power to the benchmark autoregressive process for implied volatility and, if so, what is the most appropriate description of the process? This issue is investigated at both weekly and monthly frequencies since the resulting relationship is used as a basis for several of the subsequent tests. Second, is there evidence that volatility is significantly higher in bear than in bull markets. Third, if so, does this relationship survive once the role of high-frequency, week-by-week movements is allowed for?

Nesting directionality in a more general relationship

The procedure for investigating the general relationship between implied volatility (IVB) and market movements was to add the percentage change of the bond yield (ΔR)[122] during the observation interval to the autoregressive equation for implied volatility. The change was split between positive (ΔR^+) and zero or negative observations (ΔR^-) i.e.

$$(1)\quad IVB_t = \alpha + \beta\, IVB_{t-1} + \gamma^+ \Delta R_t^+ + \gamma^- \Delta R_t^-$$

The various types of link between implied volatility and market movements impose restrictions on this general form (Table AIII.1).

In Case 1 (no relationship), proximate market movements have no effect on volatility. In Case 2 (symmetry or adaptivity), both positive and negative changes raise volatility i.e. it is the absolute value of the change that matters, not its sign. In Case 3 (weak asymmetry), the impact of increases in yields (market declines) is stronger than that of reductions in yields. In Case 4 (semi-directionality), only increases in bond yields have an effect. Finally, in Case 5 (directionality), declines in yields actually reduce volatility by as much as increments raise it. Other configurations are of course possible but the ones just mentioned turned out to be the more relevant ones.

[122] Approximated by the first difference in the log of the yield.

Table AIII.1

Types of relationship between volatility and market movements

Case 1	$\gamma^+ = \gamma^- = 0$	no relationship
Case 2	$\gamma^+ = -\gamma^- > 0$	symmetry (adaptivity)
Case 3	$\gamma^+ > 0, \gamma^- < 0, \gamma^+ > -\gamma$	weak asymmetry
Case 4	$\gamma^+ > 0, \gamma^- = 0$	semi-directionality
Case 5	$\gamma^+ = \gamma^- > 0$	directionality

Table AIII.2 summarises the results of the regressions performed on weekly data; shadings signal the preferred configuration, all things considered. The regressions indicate that in most countries the preferred relationship takes the form of semi-directionality. In Japan, by contrast, it is clearly the absolute value of the change in bond yields that matters; in Sweden and Spain, this model is only marginally preferred to that specifying that only increments in bond yields are relevant. In the United States, proximate changes in bond yields have no effect. The result is similar for Canada, although increments in bond yields are almost statistically significant at the 10% level.

An interesting feature of the findings is that pure directionality and, less often, the absolute change in the yield are significant *when taken in isolation*. In the absence of an encompassing strategy, therefore, it is quite easy to end the search at the wrong point. In several cases directionality and adaptivity appear to be acceptable representations simply because in the general specification yield declines are strongly statistically insignificant. These models, that is, derive all of their explanatory power from the impact of yield increases.

The preferred equations are reported in Table AIII.3. The incremental explanatory power in relation to the benchmark AR(1) autoregression is typically between 0.02 and 0.07. In all cases, the new variable helps to reduce the point-estimate of persistence, usually by some 3 to 5 percentage points.

The results at the monthly frequency using month-end data are broadly consistent with those at the weekly frequency (Table AIII.4).[123]

[123] The results of regressions based on month-average numbers did not differ much from those based on month-end statistics.

Table AIII.2

Volatility and market movements: regression results

Weekly data

	General specification		Individual variables[1]				Tests (% probability)[2]			
	γ^+	γ^-	ΔR^+	$-\Delta R^-$	$\lvert\Delta R\rvert$	ΔR	$\gamma^-=0$	$\gamma^+=0$	$\gamma^+=-\gamma^-$	$\gamma^+=\gamma^-$
US	0.10	0.03	0.08	0.02	0.06	0.03	71.4	21.0	33.2	33.9
	(0.08)	(0.07)	(0.07)	(0.06)	(0.07)	(0.04)				
JP	0.39***	0.35***	0.25**	0.20**	0.37***	0.02	0.0***	0.0***	74.5	0.0***
	(0.13)	(0.09)	(0.11)	(0.08)	(0.09)	(0.07)				
DE	0.50***	0.08	0.48***	0.29**	0.43***	0.29***	54.4	0.4***	0.1***	3.5**
	(0.17)	(0.12)	(0.15)	(0.11)	(0.17)	(0.09)				
FR	0.66***	0.05	0.67***	0.40***	0.45***	0.40***	56.3	0.0***	0.0***	0.0***
	(0.12)	(0.08)	(0.11)	(0.10)	(0.12)	(0.07)				
UK	0.18***	0.05	0.16***	0.02	0.12**	0.07*	50.2	0.2***	10.3	4.8**
	(0.06)	(0.08)	(0.05)	(0.08)	(0.05)	(0.04)				
IT	0.44***	0.01	0.43***	0.18*	0.25**	0.23***	93.5	0.2***	0.0***	2.4**
	(0.14)	(0.08)	(0.13)	(0.09)	(0.11)	(0.07)				
CA	0.12	0.03	0.11[3]	−0.04	0.08	0.05	70.3	12.6	20.3	27.4
	(0.08)	(0.07)	(0.07)	(0.06)	(0.07)	(0.04)				
BE	0.31**	0.01	0.30**	0.13	0.19*	0.16**	90.9	2.6**	4.9**	12.1
	(0.14)	(0.12)	(0.13)	(0.12)	(0.10)	(0.08)				
NL	0.36***	0.10	0.32***	0.08	0.27***	0.16**	21.9	0.1***	1.1**	0.6***
	(0.11)	(0.08)	(0.10)	(0.08)	(0.10)	(0.06)				
ES	0.60***	0.46*	0.44***	0.18	0.54***[4]	0.13	9.8*	0.0***	53.1	0.3***
	(0.09)	(0.28)	(0.07)	(0.27)	(0.15)	(0.12)				
DK	0.47***	−0.11	0.50***	−0.55***	0.46***	0.36***	53.2	0.0***	0.0***	11.5
	(0.09)	(0.17)	(0.08)	(0.20)	(0.09)	(0.06)				
SE	0.34***	0.23	0.29***	−0.08	0.34***[4]	0.18***	13.1	0.0***	46.8	0.4***
	(0.08)	(0.15)	(0.06)	(0.14)	(0.07)	(0.05)				
AU	0.62***	0.04	0.61***	0.25**	0.42***	0.34***	74.3	0.0***	0.0***	0.4***
	(0.14)	(0.11)	(0.12)	(0.10)	(0.15)	(0.08)				

[1] Coefficients for each of the variables shown taken in turn in an AR(1) regression for implied volatility.
[2] Marginal significance level for the hypothesis shown (F-test). [3] Marginal significance level equal to 10.4%.
[4] Adaptivity preferred to semi-directionality only on the basis of the incremental adjusted $\bar{R}^2$ (0.04 and 0.003 for Spain and Sweden respectively).

The main differences are that for Canada now the absolute value of the change in the yield becomes significant and that in the case of the United Kingdom the same model is preferred to that postulating an asymmetric relationship. When it comes to choosing between weekly and monthly findings as a representation of the underlying relationship, the finer infor-

Table AIII.3

Volatility and market movements: preferred regression

Weekly data

	Const.	ΔR_t^+	ΔR_t^-	$\lvert\Delta R_t\rvert$	IVB_{t-1}	$\bar{R}^2$	$\Delta\bar{R}^2$ [1]	DW	ΔIVB_{t-1} [2]
US	1.64**				0.90***	0.81	–	2.07	–
	(0.65)				(0.04)				
JP	1.13*			0.37***	0.89***	0.89	0.02	2.41	–0.05
	(0.58)			(0.09)	(0.04)				
DE	0.65**	0.48***			0.92***	0.95	0.02	2.24	–0.04
	(0.25)	(0.15)			(0.03)				
FR	1.54**	0.67***			0.86***	0.88	0.07	2.47	–0.04
	(0.62)	(0.11)			(0.05)				
UK	0.74*	0.16***			0.94***	0.92	0.004	2.33	–0.01
	(0.38)	(0.05)			(0.03)				
IT	2.71***	0.43***			0.81***	0.79	0.07	2.20	–0.03
	(0.65)	(0.13)			(0.04)				
CA	0.89**				0.95***	0.90	–	1.82	–
	(0.42)				(0.03)				
BE	0.87*	0.30**			0.91***	0.91	0.01	2.26	–0.03
	(0.47)	(0.13)			(0.04)				
NL	0.60**	0.32***			0.93***	0.95	0.01	2.19	–0.03
	(0.28)	(0.10)			(0.03)				
ES	3.81***			0.55***	0.71***	0.67	0.09	2.54	–0.05
	(1.36)			(0.15)	(0.09)				
DK	1.23**	0.50***			0.89***	0.91	0.08	1.82	–0.03
	(0.57)	(0.08)			(0.04)				
SE	1.04			0.34***	0.91***	0.91	0.03	2.06	–0.03
	(0.74)			(0.07)	(0.04)				
AU	3.07**	0.61***			0.81***	0.85	0.07	1.76	–0.07
	(1.25)	(0.12)			(0.07)				

[1] Incremental $\bar{R}^2$ compared with simple AR(1) regression. [2] Change in the autoregressive coefficient compared with simple AR(1) regression.

mation contained in the weekly data sways the balance in favour of the high frequency regressions.[124]

Volatility in bear and bull markets

The procedure to test econometrically whether bond yield volatility is higher in bear tan in bull markets consisted of two steps. First, the period

[124] But, of course, the best representation at each frequency was used as basis for the corresponding tests.

Table AIII.4

Volatility and market movements: preferred regression

Month-end data

	Const.	ΔR_t^+	ΔR_t^-	$\lvert\Delta R_t\rvert$	IVB_{t-1}	$\bar{R}^2$	$\Delta\bar{R}^2$ [1]	DW	ΔIVB_{t-1} [2]
US	4.63**				0.72***	0.50	–	1.75	–
	(2.00)				(0.12)				
JP	0.95			0.51***	0.77***	0.85	0.20	1.93	–0.05
	(1.22)			(0.09)	(0.07)				
DE	1.66	0.77***	0.26**		0.75***	0.88	0.09	2.41	–0.12
	(0.86)	(0.25)	(0.12)		(0.08)				
FR	4.04**	0.53**			0.67***	0.78	0.10	2.41	–0.16
	(1.80)	(0.24)			(0.14)				
UK	1.29			0.24**	0.86***	0.73	0.05	2.29	–0.03
	(1.04)			(0.09)	(0.07)				
IT	5.68***	0.36**			0.62***	0.62	0.10	1.84	–0.08
	(1.32)	(0.13)			(0.08)				
CA	1.78			0.31***	0.83***	0.74	0.10	1.40[3]	–0.02
	(1.44)			(0.08)	(0.08)				
BE	1.86**				0.85***	0.77	–	2.00	–
	(0.84)				(0.08)				
NL	2.33*	0.35*			0.77***	0.78	0.03	2.30	–0.08
	(1.29)	(0.20)			(0.14)				
ES	5.85**	0.42***			0.60***	0.49	0.16	1.76	0.01
	(2.19)	(0.12)			(0.14)				
DK	4.70***	0.58***			0.62***	0.79	0.33	2.53	–0.11
	(1.49)	(0.14)			(0.09)				
SE	3.23	0.65***			0.73***	0.83	0.36	2.28	0.01
	(2.08)	(0.17)			(0.10)				
AU	7.58**	0.75***	0.24**		0.48***	0.76	0.53	2.41	–0.01
	(2.85)	(0.18)	(0.10)		(0.14)				

[1] Incremental $\bar{R}^2$ compared with simple AR(1) regression. [2] Change in the autoregressive coefficient compared with simple AR(1) regression. [3] A second (statistically significant) autoregressive lag eliminates the signs of serial correlation; the coefficient for ΔR^+ and corresponding standard error are almost identical to the AR(1) process (0.31 and 0.07 respectively).

under consideration was divided into bull/bear markets. A bull (bear) market was defined as a period of *protracted* falls (increases) in bond yields. The choice was done simply by inspecting the graphs and picking turning points on the basis of *daily* data. Second, the statistical significance of differences between the average level of volatility in bear and bull markets was tested as a set of restrictions on a AR(1) regression where the constant and autoregressive parameters were allowed to vary as between the two types of market.

Inspection of the movements in bond yields indicated a sequence of bull-bear-bull markets between September 1992 and May 1995 in all countries (Table AIII.5). After tending to fall, yields began to rise almost everywhere in January-early February 1994. The main exception was the United States, where the onset of the bear market was in October 1993. The beginning of the subsequent bull market is far less coincident across countries, stretching from August 1994 in Sweden to March 1995 in Italy and Spain.

The point estimates of average volatility calculated over periods of falling and rising yields support the claim that volatility is indeed higher in bear markets (same table). The gap typically ranges between two to five percentage points. In percentage terms, it is anywhere between 10 and 60% of the level of volatility in the bull market. The difference is smallest in the United States and largest in core ERM countries and Japan.

A natural way of testing whether the differences are also *statistically* significant is to do a difference-in-means test.[125] In this case, the results indicate that in *all* countries the probability that volatilities are indeed equal in bear and bull markets is extremely low (same table).

The foregoing test, however, ignores the fact that volatilities are strongly autocorrelated over time. Is the claim still valid once persistence is taken into account? To test this, consider as the starting point an AR(1) regression where both the constant and the autoregressive parameters are split between bull (L) and bear (R) markets

(2) $IVB_t = \alpha_L + \alpha_0 \, DUM_{R,t} + \beta_L IVB_{L,t-1} \; \beta_R \, IVB_{R,t-1}$

where $DUM = 1$ if bear market
where $DUM = 0$ otherwise

then $\alpha_R = \alpha_L + \alpha_0$

The estimates of (unconditional) volatility for bear and bull markets implied by this equation are, respectively[126]

[125] This was implemented by running implied volatility on a constant and an additive dummy flagging bear markets and testing for the statistical significance of the dummy (see below).

[126] To see this, simply set $IVB_t = IVB_{t-1} = IVB$ in bear and bull markets and solve in terms of the underlying parameters. This procedure, of course, assumes that volatility is stationary.

Table AIII.5

Average volatility in bull and bear markets

	Trough	Peak	Average volatility[1]				
			bear	bull	difference		test (% probability)[2]
					absolute	as a % of bull	
US	17.10.93	7.11.94	17.4	15.8	1.6	10	0.00***
JP	28.12.93	1.09.94	21.9	14.9	7.0	47	0.00***
DE	13.01.94	8.01.95	15.4	9.3	6.1	66	0.00***
FR	13.01.94	9.01.95	17.8	12.6	5.2	41	0.00***
UK	27.12.93	20.09.94	19.2	14.0	5.2	37	0.00***
IT	3.02.94	20.03.95	17.9	15.4	2.5	16	0.00***
CA	30.01.94	22.01.95	18.5	16.1	2.4	15	0.00***
BE	3.01.94	12.09.94	15.7	10.8	4.9	45	0.00***
NL	3.01.94	9.01.95	15.1	9.6	5.5	57	0.00***
ES	1.02.94	30.03.95	17.9	16.0	1.9	12	0.01***
DK	19.01.94	12.09.94	18.5	13.2	5.3	40	0.00***
SE	1.02.94	17.08.94	22.5	20.2	2.3	11	1.39**
AU	31.01.94	21.11.94	21.9	17.3	4.6	27	0.00***

[1] Measured as the annualised standard deviation of percentage changes in yields (approximated by the first difference in the logarithms). [2] Test that the difference in volatility in the two types of market is statistically significant, not taking into account the autocorrelation structure.

(3) $\hat{IVB}_L = \alpha_L / 1-\beta_L$

(3') $\hat{IVB}_R = \alpha_R / 1-\beta_R = (\alpha_L + \alpha_0) / 1-\beta_R$

One way of testing the hypothesis that the two are equal is to equate (3) and (3'). The problem with this approach is that when the autoregressive parameter is close to one, the linear approximation may not be accurate enough. The second, preferred possibility is to test two restrictions

$$\alpha_0 = 0 \Rightarrow \alpha_L = \alpha_R$$
$$\beta_L = \beta_R$$

The formal tests suggest that the differences in volatility are often statistically significant, at least at the 10% level, but the results are not unambiguous (Table AIII.6). The joint test rejects the equality of coefficients only in four out of thirteen cases (Japan, France, United Kingdom and Denmark) although it is close to rejection in at least another five

Table AIII.6

Volatility in bull and bear markets: regression results

Weekly data

	Tests (% probability)[1]			DUM		Derived volatilities[2]	
	$\beta_L=\beta_R$ $\alpha_L=\alpha_R$	$\beta_L=\beta_R$	$\alpha_L=\alpha_R$ given β[3]	coeff.	SE	bull	bear
US	55.9	28.2	80.0	0.04	0.14	16.3	16.7
JP	7.8**[4]	3.5**	61.2	0.23	0.45	16.8	19.8
DE	11.5	49.7	4.8**	0.84	0.42	9.3	15.6
FR	3.0**	37.7	1.0***	1.23	0.47	12.5	18.0
UK	1.9**	46.5	0.5***	0.80	0.28	13.7	20.3
IT	11.8	17.9	9.8*	0.44	0.26	15.7	17.9
CA	29.4	11.8	45.6	0.12	0.16	16.1	18.1
BE	14.7	45.5	5.7*	0.57	0.29	10.9	16.1
NL	16.0	67.8	21.4	0.42	0.33	9.9	15.0
ES	12.0	7.1*	15.6	0.58	1.42	15.9	18.0
DK	2.8**	21.9	4.9**	0.97	0.48	12.7	18.6
SE	11.3	30.9	6.0*	0.86	0.45	17.1	27.4
AU	10.9[5]	16.5	33.53	0.50	0.52	17.9	20.7

[1] Percentage marginal probability of incorrect rejection of the null-hypothesis shown (F-test). [2] Volatility point-estimates derived from the parameters assuming the same autoregressive coefficient in bear and bull markets. [3] Tests for a zero dummy conditional on a common autoregressive parameter. [4] Perverse result: volatility higher in bull market. The result is sensitive to an autoregressive coefficient almost equal to one in the bull market. [5] Significant according to the asymplotic test (10%).

(Germany, Italy, Spain, Sweden and Australia). However, conditional on the equality between autoregressive coefficients, which is generally accepted, the tests indicate that the two volatilities are statistically different in a majority of countries. Exceptions include the United States, Canada, Japan, the Netherlands and Spain. For the Netherlands the result is rather puzzling, given the large gap between the standard deviations in the two samples (Table AIII.5).

Market vs. week-by-week effects

Given that the previous evidence indicated the existence of some form of week-by-week directionality in most markets, it is natural to ask whether higher volatility is a property of medium-term market movements or of short-term ones. In order to answer this question the equation splitting the parameters on the basis of bull and bear markets was augmented with the week-by-week term found relevant for most markets. The two hypotheses, that is, were allowed to compete against each other.[127]

The results tended to favour clearly the week-by-week effect. *All* the week-by-week effects remained significant while *none* of the tests for the equality of coefficients was rejected (not reported).[128]

On balance, we interpret the evidence as suggesting that *over the period under consideration* volatility is statistically significantly higher in bear than in bull markets in most countries. The clearest exceptions are the United States, Canada and Japan. Such higher volatility generally appears to be the result of a greater sensitivity to *short-run* increases than to decreases in yields (the week-by-week asymmetric effect). Finally, the week-by-week asymmetric effect is generally present even in cases where the difference between bear and bull market volatility is not statistically evident at stringent levels of significance.

[127] This was of course not done for the United States and Canada, where there was no evidence of the impact of proximate market movements.

[128] Actually, in the case of Italy the test was significant at the 10% level but the implied point estimate indicated lower, not higher, residual volatility in bear markets.

Annex IV – Estimating the money market volatility process and its link with bond yield volatility

In the main text it was argued that the OLS estimates of the relationship between implied bond yield volatility and historical money market volatility measured on a weekly basis were likely to represent a *lower bound* to the true strength of the link. This annex explains why and describes the technique used to get a view of the adjustment required to correct this bias.[129] It concludes that the adjustment varies considerably across countries but is typically in the region of two to four times the OLS point estimate.

What follows is rather technical but the main ideas are actually quite simple. The basic point is that measuring volatility over very short horizons necessarily relies on few observations. As a result, the volatility so measured is likely to vary considerably more than underlying or "true" volatility i.e., in economic language, to be a "noisy" estimate of it. Agents are bound to discount these movements when making decisions. The "noise", therefore, will tend to cloud the underlying relationship. It is, however, possible to pierce this veil by making some assumptions about the underlying volatility process and estimating the corresponding parameters making use of such information. The assumptions impose sufficient structure on the observed movements to distinguish the "noise" from the true "signal". The same information can then be used to correct the raw estimates when considering the relationship between implied bond yield volatility and money market volatility.

The money market volatility process

The main hypothesis required is that "true" volatility (V^*) follows an autoregressive process of some order. In what follows, it will be assumed that one autoregressive lag is a good approximation. This hypothesis is in line with most of the econometric work on volatility. Observed weekly volatility (V) measures underlying volatility with some error μ_t. The foregoing assumptions, together with additional restrictions on probability distributions, are captured in equations (1) and (2) below:[130]

[129] We are grateful to Stefan Gerlach for suggesting this possibility.
[130] Equation (1) is analogous to the stochastic model of volatility. See Hull and White (1987) and Heynen and Kat (1994).

$$(1) \quad V^*_t = \alpha + \beta V^*_{t-1} + \omega_t \qquad \omega_t \rightarrow N(0, \sigma^2_\omega)$$

$$(2) \quad V_t = V^*_t + \mu_t \qquad \mu_t \rightarrow N(0, \sigma^2_\mu)$$

More precisely, the error in the volatility process (ω_t) and in the measurement of true volatility (μ_t) are assumed to be normally distributed, serially and mutually uncorrelated processes with zero mean and constant variance and uncorrelated with the variables V^*_{t-1} and V^*_t. Given (1) and (2), the process for historical volatility, the only observable variable, can be written as

$$(3) \quad V_t = \alpha + \beta V_{t-1} + \phi_t$$

where

$$(3') \quad \phi_t = \omega_t + \mu_t - \beta\mu_{t-1}$$

It is a well known econometric result that estimating (3) by OLS yields biased parameters regardless of the sample size. The reason is that the regressor V_{t-1} and the composite error term ϕ are correlated via μ_{t-1}, given the relationship between historical and realised volatility (equation (2)). The estimate of the autoregressive coefficient β will be downward biased because the correlation between V_{t-1} and the composite error term ϕ is negative. The link between V_{t-1} and V_t is attenuated because relatively high values of V_{t-1} are associated with relatively low values of ϕ_t, which tend to drag V_t down. OLS, by contrast, assumes that V_{t-1} and ϕ_t are uncorrelated and cannot therefore disentangle the two effects.

The most common way of avoiding the bias is to estimate (3) through some instrumental variable (IV) technique. The trick is to retain only those movements of V_{t-1} that are uncorrelated with ϕ by regressing V_{t-1} on some other variable (instrument) uncorrelated with ϕ.[131] The problem is that the resulting estimates will generally be very imprecise. For example, the model restrictions do imply that further lagged values of V_t (i.e. V_{t-2}, V_{t-3} etc.) are uncorrelated with ϕ. But precisely owing to measurement error their measured correlation with V_t is bound to be very low.

[131] See Pagan (1984) and Pagan and Ullah (1988), who advocate this technique precisely in the present context, where volatilities are measured with error.

The alternative proposed here is fully to use the structure imposed by the model. In particular, (3') indicates that the composite disturbance ϕ will follow a moving average process, MA(1), estimated as

$$(3'') \quad \phi_t = \varepsilon_t - \delta\varepsilon_{t-1}$$

so that (3) is ARMA (1, 1).[132] This can be estimated accordingly. Note that because of the presence of the error term in the true volatility process, ω_t, the *estimated* MA coefficient in (3"), δ, will not generally result equal to β. In other words, there is measurement error here too. However, the autoregressive coefficient of interest, β, will be correctly estimated. Intuitively, the ARMA estimation techniques treats ε_t as an additional regressor, duly taking into account any correlation with V_{t-1}.

The further advantage of this procedure is that it yields estimates of the variances of the two errors in the model, σ^2_ω and σ^2_μ. The variances of the errors can be calculated by writing the variance and covariance of the composite error term, ϕ_t, in terms of the model parameters, equation (3'), and of the error term in the estimated ARMA representation, ε_t or equation (3"). Specifically, we have:

$$(4) \quad var(\phi) \quad \Rightarrow \quad \sigma^2_\omega + (1 + \beta^2)\sigma^2_\mu = \sigma^2_\varepsilon(1 + \delta^2)$$

$$(5) \quad cov(\phi_t, \phi_{t-1}) \quad \Rightarrow \quad -\beta\sigma^2_\mu = -\delta\sigma^2_\varepsilon$$

This yields two equations in two unknowns σ^2_μ, σ^2_ω, given that the regression provides estimates of β, δ and σ^2_ε.

The ratio of the variance of the error term in the true volatility process to that in the measurement error in true volatility, $\sigma^2_\omega / \sigma^2_\mu$, is of some interest. The lower this ratio, the more likely it is that the estimated MA parameter δ will equal β. But the variance of the measurement error, σ^2_μ, is more important. This parameter holds the key to calculating the bias in the OLS estimates and hence to adjusting them. A standard econometric result is that this bias depends on the ratio of the variance of the measurement error to that of the "true" variable, σ^2_{V*}. More precisely, letting $\hat{\beta}$ stand for the OLS estimate, we have

[132] More generally, if V^*_t is AR(i), then V_t is ARMA (i,i).

$$(6)\quad \hat{\beta} = \frac{1}{(1 + \sigma^2_\mu / \sigma^2_{V*})}\beta$$

And σ^2_V can easily be calculated. Given (2),

$$(7)\quad \sigma^2_{V*} = \sigma^2_V - \sigma^2_{\mu*}$$

with σ^2_V being the variance of observed volatility, V_t, over the relevant sample period.

Table AIV.1 compares the estimates of AR(1) process for money market volatility derived from OLS, IV and ARMA estimation techniques. The basic data are non-overlapping money market historical volatilities measured on a weekly basis.[133] One sample period always corresponds to that over which the relationship between implied bond yield volatility and money market volatility was assessed; in two cases a different sample was also used. A number of findings stand out.

The results confirm the downward bias in OLS estimates of β. These are typically considerably lower than either the IV or ARMA ones. The Hausman test clearly signals the existence of measurement error.

The IV and ARMA estimates are often quite close, though the difference may be sizable in certain cases. As expected, however, the standard error of the autoregressive coefficient derived from the ARMA model is much lower than that obtained through the IV technique. The efficiency gains are apparent.

The ARMA point-estimates of the autoregressive parameter vary considerably across countries. They tend to be comparatively low in most continental Europe; at .35, the parameter is lowest in Germany. At the other end of the spectrum, the coefficient is essentially unity in the United States, Sweden and Denmark and only slightly lower in the United Kingdom.

Table AIV.2 reports the *theoretical* adjustment factors for the OLS estimates based on the implied estimates for the relevant variances, together with other interesting statistics. The adjustment factors range from 1.0 (no adjustment required) in the case of Australia and Sweden (when the relationship is estimated since August 1992) to 5 or over in those of the

[133] The variable is the same as that used in the text: the percentage change in the implied 3-month rate three-month forward. The volatility is measured around an imposed zero mean over five business days.

Table AIV.1

Estimating money market volatility: basic results[1]

	β			δ	$\bar{R}^2$		HAUSMAN	β=δ
	OLS	IV	ARMA	ARMA	OLS	ARMA		
US	0.19** (0.08)	1.07** (0.47)	1.00*** (0.01)	1.02*** (0.02)	0.03	0.21	0.00***	44.6
JP	0.25*** (0.08)	0.95*** (0.33)	0.89*** (0.08)	0.72*** (0.13)	0.06	0.10	0.00***	1.2**
DE	0.13 (0.08)	0.15 (0.54)	0.35 (0.42)	0.23 (0.44)	0.01	0.01	0.00***	17.1
FR	0.40*** (0.08)	0.68*** (0.19)	0.64*** (0.14)	0.29 (0.18)	0.16	0.17	0.00***	0.0***
UK	0.44*** (0.08)	0.78*** (0.18)	0.96*** (0.02)	0.94*** (0.03)	0.19	0.29	0.00***	43.0
IT	0.23*** (0.08)	0.57 (0.34)	0.50* (0.28)	0.28 (0.31)	0.05	0.05	0.00***	0.9***
CA	0.22*** (0.08)	0.89*** (0.30)	0.94*** (0.05)	0.82*** (0.08)	0.04	0.13	0.00***	1.0***
BE	0.42*** (0.08)	0.64*** (0.18)	0.79*** (0.10)	0.50*** (0.15)	0.17	0.18	0.00***	0.0***
NL	0.39*** (0.08)	0.66*** (0.19)	0.74*** (0.12)	0.43*** (0.16)	0.15	0.16	0.00***	0.0***
ES	0.36*** (0.10)	0.69** (0.27)	0.68*** (0.19)	0.35 (0.24)	0.09	0.10	0.00***	0.2***
DK	0.23* (0.12)	0.89* (0.49)	0.99*** (0.03)	0.98*** (0.02)	0.04	0.19	0.00***	69.5
SE(1)	0.31*** (0.12)	0.67** (0.41)	1.03*** (0.02)	1.04*** (0.09)	0.08	0.23	0.00***	80.8
(2)[2]	0.62*** (0.07)	0.62*** (0.11)	0.62*** (0.11)	0.00 (0.13)	0.38	0.37	0.00***	0.0***
AU(1)	0.45*** (0.11)	0.70 (0.26)	0.45*** (0.22)	−0.01 (0.25)	0.20	0.19	0.00***	0.1***
(2)[2]	0.48*** (0.07)	0.55*** (0.16)	0.60*** (0.14)	0.16 (0.17)	0.22	0.22	0.00***	0.0***

[1] Weekly data. Unless otherwise stated, the sample period is the longest one applied to the regressions for implied bond yield volatility for each country. [2] Estimated over 31.08.92 to 22.05.95.

United States and Denmark. The corrections are typically within the 2.0 to 4.0 range. Note that for Sweden the correction is as high as 4.6 when the regression is estimated only since 1994, the same period for which the link between implied bond yield volatility and money market volatility is assessed. The large difference appears to arise from the extreme values

Table AIV.2
Additional information about relative variances

	Mean(V)	SD(V)	SD(μ)/ Mean(V)	SD(ω)/ Mean(V)	Var(ω)/ Var(μ)/	Var(μ)/ Var(V*)	Adj. factor
US	26.0	14.1	0.49	0.01	0.00	4.18	5.18
JP	58.7	30.7	0.45	0.13	0.08	2.70	3.70
DE	16.8	8.2	0.40	0.26	0.44	2.03	3.03
FR	35.1	20.0	0.35	0.35	0.45	0.60	1.60
UK	28.3	19.6	0.57	0.03	0.00	2.23	3.23
IT	33.4	25.2	0.55	0.45	0.68	1.14	2.14
CA	55.8	30.1	0.47	0.09	0.03	3.14	4.14
BE	33.0	26.7	0.58	0.34	0.35	1.06	2.06
NL	14.1	9.8	0.49	0.34	0.49	0.95	1.95
ES	29.4	20.1	0.47	0.39	0.69	0.90	1.90
DK	34.7	19.1	0.49	0.01	0.00	4.04	5.04
SE (1)	35.2	25.6	0.64	0.02	0.00	3.60	4.60
(2)	51.2	53.4	0.40	0.82	425.21	0.00	1.00
AU (1)	47.3	25.4	–	0.89	–	–	1.00
(2)	41.8	23.9	0.26	0.41	2.41	0.27	1.27

reached by money market volatility during the ERM crisis in September 1992 (over 400%). This suggests that the procedure results are quite sensitive to such episodes. The implication is that estimated measurement errors in more tranquil periods may be quite different. Presumably, however, the appropriate adjustment factor is the one corresponding to the estimation period under consideration, irrespective of how special the given period is.

There appears to be little relationship between the adjustment factor and either the mean or standard deviation of historical volatility (same table). By contrast, the adjustment factor rises with the estimate of the "true" autoregressive coefficient derived from the ARMA procedure.[134] The ratio of the standard deviation of the measurement error (μ_t) to the mean of historical weekly volatility generally ranges between 40 and 60%. The ratio of the variance of the error in the underlying volatility process (ε_t) to that of the measurement error ranges much more widely and

[134] Germany is the notable exception to this pattern: the adjustment factor is comparatively large given the point-estimate of the AR(1) coefficient. One may be tempted to attribute this to the lack of precision in the ARMA estimates, which are not statistically significant. At the same time, the evidence below indicates that the ARMA estimates should not be discounted so readily.

consistently with the evidence on the acceptance or rejection of the equality between the estimated autoregressive and moving average coefficients.

How reliable are the above estimates and corresponding adjustment factors? One way to assess this is to compare the "theoretical" adjustment to the OLS estimates with the coefficients derived from the original OLS estimates. For example, given the ARMA estimate of β and of the relevant variances, how accurate is the model in predicting the OLS estimation results? Accuracy suggests that the estimates of the variances and of β are reliable. A second possibility is to perform the whole exercise on another volatility series.

Comparing the theoretical or predicted OLS estimates of the autoregressive coefficient with the original OLS results gives credence to the whole approach. Quite remarkably, the point-estimates predicted on the basis of the adjustment factor and the ARMA autoregressive coefficient are extremely close to the actual OLS ones (Table AIV.3). By the same token, applying the adjustment factors to the original OLS coefficients gives point-estimates which are close to the ARMA estimates; non-linearities account for the percentage point differences. The most notable, and puzzling, exception is the United Kingdom, for which the OLS estimate is missed by .14 percentage points. The case is puzzling given the precision in the estimation of the AR and MA coefficients of the process.

As a further check, the same procedure was applied to weekly bond yield historical volatilities. The results were generally similar to those already found in the case of money market volatility (Table AIV.4). Moreover, there was a broad correspondence between the point-estimates of the autoregressive coefficient of the historical volatility process and those of *implied* volatility (same table). This is quite comforting. Although there is no good reason to expect implied and historical volatility to exhibit the same degree of persistence, the two processes should not diverge markedly either.

Adjusting the link with implied bond yield volatility

We are now in a position to adjust the OLS estimates of the relationship between implied bond yield volatility and historical money market volatility. Consider one such regression of the form

$$(8) \quad IVB_t = \alpha_0 + \gamma V^*_t + \delta Z_t + \eta_t$$

Table AIV.3

Relationship between actual, predicted and adjusted OLS ß coefficients

	Predicted[1]	Actual minus predicted	Adjusted[2]	ARMA minus adjusted
US	0.19	–0.01	0.97	0.03
JP	0.24	0.01	0.93	–0.04
DE	0.12	0.01	0.39	–0.04
FR	0.40	0.00	0.64	0.00
UK	0.30	0.14	1.42	–0.46
IT	0.24	–0.01	0.49	0.01
CA	0.23	–0.01	0.91	0.03
BE	0.38	0.04	0.85	–0.06
NL	0.38	0.01	0.76	–0.02
ES	0.36	0.00	0.69	–0.01
DK	0.20	0.03	1.17	–0.18
SE (1)	0.22	0.09	1.44	–0.41
(2)	0.62	0.00	0.62	0.00
AU (1)	0.31	–	0.31	0.14
(2)	0.55	–0.07	0.60	0.00

[1] Predicted on the basis of the ARMA estimation results. [2] Adjusted actual OLS coefficient on the basis of the ARMA estimation results.

where Z_t is an additional regressor e.g. lagged bond yield volatility itself.[135] In terms of historical volatility this can be written as

$$(9) \quad IVB_t = \alpha_0 + \gamma V_t + \delta Z_t + \xi_t$$

where

$$(10) \quad \xi_t = \eta_t - \gamma \mu_t$$

It can be shown that when only V_t^* is measured with error the bias of $\hat{\gamma}$ is exactly the same as the bias of $\hat{\beta}$ in the money market volatility regression (e.g. Greene (1993))

$$(11) \quad \hat{\gamma} = \frac{1}{(1 + \sigma^2_\mu / \sigma^2_{v*})} \gamma$$

[135] The following results hold regardless of the number of regressors Z_t.

Table AIV.4

Estimating bond yield volatility: summary of results[1]

	β (implied volatility)	β (historical volatility)					Adj. factor
		IV	AR	OLS			
				actual	predicted	difference	
US	0.90	1.14*	0.90***	0.09	0.14	–0.05	6.47
JP	0.94	0.92***	0.86***	0.48***	0.49	–0.01	1.78
DE	0.96	0.98***	0.95***	0.42***	0.41	0.01	2.30
FR	0.90	0.76***	0.83***	0.57***	0.55	0.02	1.51
UK	0.96	0.82***	0.94***	0.31***	0.31	0.00	3.00
IT	0.84	0.87**	0.80***	0.24***	0.23	0.01	3.48
CA	0.95	0.78***	0.87***	0.31***	0.29	0.03	3.01
BE	0.94	0.82***	0.86***	0.45***	0.43	0.02	2.00
NL	0.97	0.81***	0.93***	0.41***	0.40	0.01	2.34
ES	0.76	0.70**	0.76***	0.27***	0.27	0.00	2.77
DK	0.92	0.96***	0.97***	0.50***	0.45	0.05	2.16
SE	0.94	0.76***	0.96***	0.43***	0.33	0.10	2.86
AU[2]	0.88	0.71***	0.91***	0.28***	0.26	0.02	3.45

[1] Based on the annualised standard deviation of daily percentage changes (log of first differences) in bond yield volatility calculated over non-overlapping weeks (Fridays to Thursdays) around an imposed zero mean. The sample is the same as that used for the estimates of the AR(1) process for implied bond yield volatility. [2] Estimated over 31.08.92 to 22.05.95 owing to the presence of a negative MA root in the shorter sample.

In other words, the adjustment factor for $\hat{\gamma}$ has already been calculated.

The same cannot be said for the standard error, however. The correct standard error is the one that applies to the underlying regression (8).[136] Estimating it is relevant if one is interested also in the statistical significance of the identified coefficients. The variance of the estimate of γ in equation (8) can be shown to be

$$(12)\quad \mathrm{var}(\gamma) = \frac{\sigma_\eta^2}{(1 - R_a^2)\sigma_{v^*}^2 n}$$

where n is the number of observations in the regression for implied volatility and R_a^2 is the coefficient of determination (R-squared) in a

[136] The procedure below does not take account of the White adjustment for heteroskedasticity.

regression of the *underlying* money market volatility on the other regressors in the equation for implied bond yield volatility.[137] Intuitively, the larger the R-squared, the higher the correlation between money market volatility and the other explanatory variables and hence the harder it is to disentangle their influence.

It is in fact possible to estimate both σ^2_η and R^2_a given the information available. Since there is no reason to assume that η_t and the measurement error, μ_t, are correlated, from (10) we have

$$(13) \quad \sigma^2_\eta = \sigma^2_\xi - \gamma^2\sigma^2_\mu$$

Similarly, it can be shown that

$$(14) \quad R^2_a = \frac{\sigma^2_v}{\sigma^2_{v*}} R^2_b$$

where R^2_b is the R-squared of the regression of *observed*, rather than underlying, money market volatility on the other explanatory variables in the regression for implied bond yield volatility.[138, 139]

Table AIV.5 summarises the adjusted coefficients and standard errors.[140] By way of comparison, it also reproduces the original OLS results and a set derived through an instrumental variable technique.[141]

The adjusted OLS estimates now suggest that between 2 and 7% of money market volatility shows up in bond yield volatility. In the cases of Denmark and, especially, Sweden, the corrected coefficients look implausibly large. This may be due to the comparatively few observations available, which invalidate the asymptotic approximation used. This is confirmed by the fact that the adjusted standard errors cannot be

[137] See e.g. Greene (1993), p.268.

[138] In the single regressor case, $R^2_a = c^2\sigma^2_z / \sigma^2_{v*}$ and $R^2_b = c^2\sigma^2_z / \sigma^2_v$ where c is the common coefficient on Z_t in the regressions for either observed or underlying money market volatility. The result in the text follows immediately.

[139] Note that the adjusted standard errors may either be higher or lower than the original OLS ones.

[140] These refer to the simple AR(1) regressions with money market volatility as the only additional regressor.

[141] To facilitate comparison, none of the standard errors in the table has been adjusted for heteroskedasticity. The difference, however, is generally extremely small (compare with Table 9 in the main text). The main exception is the United Kingdom, where the unadjusted standard error is considerably lower.

Table AIV.5

Adjusted influence of money market volatility on implied bond yield volatility

	Coefficient			SE		
	OLS	adjusted	IV	OLS	adjusted	IV
US	0.012**	0.062	0.016	0.005	–	0.025
JP	0.004	0.015	0.080	0.005	0.014	0.082
DE	0.025**	0.074***	–0.026	0.010	0.014	0.093
FR	0.005	0.008	–0.007	0.007	0.008	0.016
UK	0.009**	0.030***	–0.009	0.004	0.006	0.010
IT	0.011	0.024***	0.021	0.007	0.007	0.018
CA	0.004	0.017***	0.010	0.003	0.004	0.008
BE	–0.003	–0.006	–0.001	0.003	0.004	0.007
NL	0.017**	0.033***	0.013	0.008	0.010	0.021
ES	0.006	0.011	–0.049*	0.008	0.012	0.029
DK	0.020**	0.103	0.067	0.010	–	0.052
SE	0.023***	0.106	0.097	0.007	–	0.064
AU (1)	0.009	–	0.041*	0.007	–	0.024

computed since they yield a negative variance: the point estimates of γ and of the variance of the measurement error appear to be too high.

The correlated standard errors now allow a statistically significant effect to be detected also in the case of Italy and Canada. For the United States, the calculation fails to occur by a very narrow margin: the derived variance of the estimate of γ is extremely small but negative. We would tend to attribute this to a slight upward bias in the estimate of γ and in the variance of the measurement error: the point estimate is within a plausible range and actually quite close to the "cleaner" influence detected on the basis of *implied* money market volatility (see Graph 18 in the main text).

The intrumental variable estimates appear to be quite unreliable in the present context. In particular, they often have the wrong sign. This appears to confirm that the quality of the instruments used is indeed quite poor.[142]

[142] Two lagged values of money market volatility were used as instruments.

Annex V – Implied bond yield volatility and money market volatility: additional econometric results

The main text reported only the results of OLS regressions of implied bond yield volatility on realised money market volatility on weekly and monthly data when the bond yield process was modelled as a simple autoregression (Table 9). This annex reports the equivalent estimates once proximate market movements are allowed as an additional influence on the bond volatility process (conditional regressions). Recall that for the United States and Canada there is no difference between the two since market movements have been found not to have any statistically significant influence.

On balance the OLS conditional regressions suggest that the relationship between implied bond yield volatility and money market volatility survives the inclusion of market movements (Table AV.1). The main difference in the cross-country pattern is that no statistically significant link is visible any more for Italy (later period, weekly), Sweden and Netherlands (monthly) while one can be detected for Spain (later period, monthly). In general, the point estimates of the parameters are of a similar order of magnitude. Moreover, given the arguments and evidence in Annex IV, there are reasons for believing that the OLS estimates underestimate the strength and, typically, the statistical significance of the link.

Table AV.1

Realised money volatility and implied bond volatility (conditional)

	Weekly			Monthly[1]		
	whole sample	earlier period	later period	whole sample	earlier period	later period
Japan	0.003	0.011	–0.012*	0.054*	0.061**	0.016
	(0.006)	(0.009)	(0.006)	(0.030)	(0.025)	(0.031)
Germany	0.017*	0.010	0.018	0.009	(0.019)	–0.011
	(0.010)	(0.008)	(0.015)	(0.065)	(0.051)	(0.122)
France	0.004	0.004	0.003	0.041**	0.028*	0.114*
	(0.004)	(0.005)	(0.008)	(0.019)	(0.013)	(0.056)
United Kingdom	0.010*	0.011**	0.014	0.004	0.016	0.067
	(0.005)	(0.005)	(0.017)	(0.019)	(0.021)	(0.171)
Italy	0.001	–0.009	0.006	0.017*	0.005	0.002
	(0.007)	(0.007)	(0.009)	(0.009)	(0.012)	(0.057)
Belgium	–0.002	0.004	–0.011	0.016[2]	0.021*	0.030
	(0.005)	(0.003)	(0.009)	(0.009)	(0.011)	(0.019)
Netherlands . . .	0.011*	0.0005	0.030***	0.013	0.021	–0.023
	(0.004)	(0.004)	(0.01)	(0.017)	(0.016)	(0.088)
Spain	0.004		–0.008	–0.031	–0.078	0.049**
	(0.006)		(0.010)	(0.033)	(0.051)	(0.018)
Denmark.			0.006			–0.011
			(0.007)			(0.040)
Sweden			0.013			0.059
			(0.009)			(0.052)
Australia			0.013**			0.026
			(0.006)			(0.032)
Japan						
(period split. .	–0.003	0.002	–0.008	0.054*	0.030**	0.079
at end-1993) . .	(0.006)	(0.006)	(0.008)	(0.030)	(0.013)	(0.048)

[1] Month-end data. [2] Probability 10.5%.

Annex VI – Econometric evidence on spillovers

In the following pages we explain the construction of the spillover map shown in the main text and present some additional results. After describing the general methodology, three issues are addressed. First, how have spillovers changed over time? Second, do spillovers survive once directionality is allowed for as an additional explanatory factor for changes in volatility in domestic markets? Third, is there specific evidence that the intensification of spillovers is a feature of more volatile markets per se?

The methodology

The basic approach is to test whether *lagged* volatility in a foreign market adds explanatory power to the volatility process in the domestic market. Contemporaneous relationships are excluded because they make it particularly difficult to distinguish between *common* shocks to the volatility processes across countries and *sequential* transmission of volatility across markets. Admittedly, the term "transmission" should be interpreted with care. It is always hazardous to interpret temporal sequences as signalling causal relationships. Nor can the approach cast light on transmission mechanisms. Nevertheless, the procedure is a useful, and quite popular, way of gathering further information about the stylised features of the relationship between volatilities across markets, a complement to, rather than a substitute for, informed interpretation.

For the purposes of the exercise, the domestic implied volatility process was specified in two ways: as a simple autoregression (AR(1)) and as an autoregression augmented, where relevant, by the corresponding proximate market movement (see Annex III). The richer alternative can be thought of as capturing any spillovers which occur over and above induced changes in the *levels* of the bond yields.

Likewise, foreign influences can in principle be defined narrowly or broadly. The narrow definition allows only for lagged *implied volatility* in the foreign market. The broad one considers in addition the lagged *proximate movement in the foreign market*. The second procedure is in the spirit of the econometric analysis of spillovers in the GARCH literature, where transmission typically occurs through the lagged innovation in the return process abroad. In order to limit the proliferation of tests, we confine the analysis to the transmission of implied volatility proper and hence adopt the narrow definition of foreign influences.

The thirteen markets in our sample make for a vast array of possible transmission patterns. To maintain the exercise manageable, pair-wise comparisons were made covering the various potential directions of influence. When more than one foreign country was found to be statistically significant in the pair-wise case, the relevant set of foreign markets were allowed to compete in an attempt to discern the dominant relationship. Of these, on statistical grounds the preferred specification retained only markets which were statistically significant, either individually or jointly. Joint statistical significance of otherwise statistically insignificant markets is a clear sign of collinearity. In this case it is not possible to distinguish the influence of the markets concerned.

An additional type of influence is what we term "mediation". Market 2 is said to "mediate" or "transmit" the influence of market 1 on market 3 if, according to the aforementioned criteria,

(i) Market 1 influences market 2 and 3 in pair-wise comparisons.
(ii) Neither market 2 nor 3 influence market 1.
(iii) Market 3 does not influence market 2.
(iv) Market 1's influence on market 3 disappears once market 2 is taken into account.

The reason for this interpretation is that a hierarchy of influence is involved, with markets 1 and 3 at opposite ends and market 2 in the middle.[143]

Drawing the spillover map

Preliminary analysis clearly indicated that it made little sense to consider the whole sample period, starting generally in August 1992, as a single one: the strength and geographical scope of spillovers was simply too variable over time. A natural date to split the sample was the tightening of monetary policy in the United States in early 1992. It was this tightening that signalled the common rise in volatility across countries.

[143] It may be tempting but profoundly misleading to think that if market 2 transmits the influence of market 1 on market 3 the whole process takes two weeks i.e. two observation intervals. As the explanation of the concept makes clear, the relevant interval remains one week. Mediation refers to different degrees of "statistical" proximity, not to a temporal sequence in real time. On the other hand, a finer observation interval could indeed reveal an underlying temporal sequence.

Table AVI.1

Unconditional spillovers, August 1992 – January 1994[1]

Domestic market	Foreign market DE	FR	UK	Own lag	$\bar{R}^2$	$\bar{R}^{2}$ [2]	ΔIVB_{t-1} [3]
UK	0.20**		–	0.89***	0.84	0.01	–0.03
	(0.11)			(0.05)			
IT	0.99**			0.50***	0.67	0.08	–0.25
	(0.48)			(0.14)			
IT		0.63*		0.61***	0.65	0.06	–0.14
		(0.32)		(0.10)			
IT	0.92**	0.57***		0.39***	0.72	0.13	–0.36
	(0.38)	(0.20)		(0.14)			
BE			0.07*	0.91***	0.87	0.01	–0.02
			(0.04)	(0.05)			
NL	0.15***			0.77***	0.83	0.03	–0.20
	(0.05)			(0.06)			

[1] There was also some evidence of France and Denmark affecting, respectively, the United States and Japan, which was discarded on a priori grounds. [2] Incremental $\bar{R}^2$ with respect to the autoregression for the domestic country. [3] Change in the persistence parameter with respect to the autoregression for the domestic country.

Tables AVI.1 and AVI.2 summarise the main linkages before and after the tightening emerging in bivariate comparisons i.e. before attempting to disentangle the dominant influences between countries. The contrast between the two sub-periods is striking. Spillovers in the earlier period were quite sparse; they were much more widespread in the later one. In particular, the almost ubiquitous presence of the United States as a "source" of volatility in the post-tightening sample stands out: only Japan escaped its influence. Note also how the United States is the only country which generally affects, but is not affected by, any other: the asymmetry is quite strong. Japan remains essentially an island.

Table AVI.3 sheds further light on the linkages in the post-US tightening period. It reports the "preferred" relationships for each country together with the pairwise results vis-à-vis those markets that survive in the transmission chain according to the criteria developed above i.e. that are either at the origin of the chain (the United States) or that mediate in the transmission.

On statistical grounds, it is possible to identify at least two groups of countries in terms of the chain of transmission. A set of "high-yielders"

Table AVI.2

Summary of pairwise unconditional spillovers, February 1994 – May 1995

Domestic country	Foreign country												
	US	JP	DE	FR	UK	IT	CA	BE	NL	ES	DK	SE	AU
US	–												
JP		–											
DE	**		–		***		**		**				*
FR	**		***	–	***		***	**	***		**		
UK	***				–								
IT	**					–					*		
CA	***	*					–						***
BE	***		***		***	**	***	–	***		***		***
NL	***				***		***		–				***
ES	**	*	**	**	***	*	***	**	***	–	***	*	***
DK	**						*				–		
SE	***				**		***	**	**			–	
AU	**												–

Note: The number of "*" indicates the level of significance.

Table AVI.3

Unconditional spillovers, February 1994 – May 1995: detailed results

	US	UK	IT	CA	NL	AU	Own lag	$\bar{R}^2$	$\Delta\bar{R}^2$	ΔIVB_{t-1}
DE	0.30** (0.12)						0.75*** (0.09)	0.85	0.02	–0.16
		0.45*** (0.15)					0.45*** (0.12)	0.86	0.03	–0.46
					0.45** (0.22)		0.44* (0.23)	0.84	0.01	–0.47
		0.45*** (0.11)			0.48** (0.20)		–0.06 (0.27)	0.87	0.04	–0.97
FR	0.38** (0.15)						0.62*** (0.11)	0.66	0.04	–0.16
		0.37*** (0.13)					0.46*** (0.15)	0.67	0.05	–0.32
					0.78*** (0.14)		0.08 (0.16)	0.74	0.11	–0.70
UK	0.30*** (0.11)	–					0.79*** (0.06)	0.90	0.01	–0.16
IT	0.18** (0.08)		–				0.79*** (0.07)	0.78	0.01	–0.09
CA	0.23*** (0.08)			–			0.82*** (0.05)	0.94	0.01	–0.14
	0.22** (0.10)			–		0.10** (0.05)	0.74*** (0.07)	0.94	0.01	–0.22
BE	0.25*** (0.08)						0.78*** (0.11)	0.84	0.02	–0.13
		0.29*** (0.60)					0.60*** (0.15)	0.85	0.03	–0.31
					0.59*** (0.13)		0.26* (0.14)	0.88	0.06	–0.65
			0.19*** (0.07)	0.18*** (0.06)	0.50*** (0.12)		0.04 (0.10)	0.89	0.08	–0.87
NL	0.28*** (0.10)				–		0.80*** (0.06)	0.90	0.01	–0.14
		0.33*** (0.09)			–		0.61*** (0.09)	0.90	0.02	–0.33
ES	0.49** (0.20)						0.48*** (0.16)	0.54	0.04	–0.22
			0.32** (0.15)	0.54*** (0.11)		0.22** (0.10)	–0.03 (0.08)	0.66	0.17	–0.73
DK	0.30** (0.12)						0.80*** (0.07)	0.85	0.02	–0.12
SE	0.25*** (0.08)						0.88*** (0.05)	0.89	0.01	–0.06
AU	0.31** (0.12)					–	0.70*** (0.10)	0.79	0.01	–0.18
	0.22** (0.10)			0.74*** (0.07)		–	0.10** (0.05)	0.94	0.17	–0.84

(Australia, Canada, Italy and Sweden), which are directly affected by the United States and have some influence on another "high-yielder" (Spain). And the set of "core" European countries, which receive the US influence via the United Kingdom (the Netherlands, Germany, France and Belgium). Puzzlingly, the Netherlands appears to help in transmitting the influence of the United Kingdom to the other countries in this group.[144] The asymmetry between the United Kingdom and The Netherlands, on the one hand, and Germany, on the other, is econometrically robust.[145]

The size of the identified effects is quite large. The spillover coefficients typically range between ¼ and ½ but are sometimes considerably higher. Moreover, the addition of the foreign market generally leads to a considerable reduction in the size of the persistence parameter: the identified influence of the proximate history of volatility in the domestic market declines. Indeed, in some cases the autoregressive parameter becomes statistically insignificant and may even have the wrong sign. Spain and, strikingly, Germany and France are cases in point.

The broad pattern of results once the relevant proximate market movements are added to the domestic volatility process is similar. Prior to February 1994 spillovers are still limited and circumscribed to Europe, with Germany playing some role (Table AVI.4). Spillovers intensify following the tightening of US monetary policy (Table AVI.5). Admittedly, they are less numerous. In particular, there is no longer evidence of a direct influence of the United States on a number of European markets. This suggests that the effects uncovered in the previous regressions were in part operating through induced changes in yields in domestic markets. Nonetheless, the preferred specifications are in general very close to the original ones, both in terms of the cross-country pattern of relationships and the size of the identified effects (Table AVI.6). A notable difference is that the influence of the Netherlands on Germany disappears: only that of the United Kingdom remains visible.

[144] If the role of the Netherlands is disregarded, it becomes impossible to disentangle the influence of Germany and the United Kingdom on either Belgium or France. On a pairwise basis, the result for Germany is very similar to that for the United Kingdom shown in Table AVI.3. When both foreign markets are included, they are individually statistically insignificant but jointly significant at the 1% (Belgium) and 5% (France) level. Moreover, the sum of the individual coefficients is close to the foreign influence in the pairwise relationships. These are obvious signs of collinearity.

[145] The coefficient (standard error) of the German market in the regressions for the United Kingdom and the Netherlands are, respectively, –0.08 (0.15) and 0.25 (0.26).

Table AVI.4

Conditional spillovers, August 1992 – January 1994[1]

	DE	FR	UK	ΔR	Own lag	$\bar{R}^2$	$\Delta\bar{R}^2$	ΔIVB_{t-1}
UK	0.21*		–	–0.04	0.89***	0.84	0.01	–0.03
	(0.11)			(0.06)	(0.05)			
IT	0.71***			0.75***	0.56***	0.78	0.04	–0.18
	(0.19)			(0.21)	(0.08)			
IT		0.34*		0.75***	0.66***	0.76	0.02	–0.08
		(0.18)		(0.28)	(0.07)			
IT[2]	0.71***	0.34***		0.64***	0.48***	0.80	0.06	–0.26
	(0.18)	(0.12)		(0.21)	(0.08)			
BE			0.07**	0.29**	0.91***	0.87	0.02	–0.02
			(0.03)	(0.13)	(0.05)			

[1] A weak effect of Belgium on Japan was ignored on a priori grounds. [2] When considered individually, Belgium is also significant at the 10% level (coefficient=0.13); it drops out, however, when either Germany or France are included in the regression.

Spillovers and high volatility

Splitting the sample in February goes some way towards answering the question of whether spillovers are stronger when volatility is high since volatility was generally higher in the more recent period. It falls short of a satisfactory answer, however, because the match is far from perfect. Since February there have been quiet periods too and spells of comparatively high volatility also occurred in 1992 and 1993 in some European countries.

In order to tackle the issue of the relationship between spillovers and the level of volatility, a specific test was therefore developed. The sample was split into two sub-samples on the basis of the level of volatility. More precisely, the top twenty observations[146] of joint highest volatility between the domestic and a foreign market were singled out. Two separate regressions were in effect then run, by allowing *all* the parameters of a single regression to differ between the two periods through appropriate dummies. Certain countries were excluded from the test as receipients of volatility i.e. the United States and Japan, owing to the inability to detect any spillovers in the previous analysis, and Denmark, Sweden and Australia, because of the much smaller number of observations available.

[146] The exercise was repeated on the basis of the top 30 observations with very similar results.

Table AVI.5

Summary of pairwise conditional spillovers, February 1994 – May 1995

Domestic country	Foreign country												
	US	JP	DE	FR	UK	IT	CA	BE	NL	ES	DK	SE	AU
US[1]	–	–	–	–	–	–	–	–	–	–	–	–	–
JP		–											
DE			–		***								*
FR				–					***				
UK	**				–								
IT						–					[2]		
CA[1]	–	–	–	–	–	–	–	–	–	–	–	–	–
BE				**		**	**	–	***		***		***
NL	**				***		**		–		*		
ES				**	***	*	**	*	**	–	**	**	**
DK											–		
SE	*											–	
AU	**												–

Note: The number of "*" indicates the level of significance.

[1] The previous results hold for the United States and Canada since, in these two countries, proximate market movements are not statistically significant. [2] Very weak evidence of a small effect from Denmark ignored.

Table AVI.6

Conditional spillovers, February 1994 – May 1995: detailed results

	US	UK	IT	CA	NL	ΔR	Own lag	$\bar{R}^2$	$\Delta\bar{R}^2$	ΔIVB_{t-1}
DE		0.26***				0.46***	0.62***	0.90	0.01	–0.26
		(0.09)				(0.16)	(0.11)			
FR					0.43***	0.55***	0.42***	0.81	0.03	–0.40
					(0.12)	(0.11)	(0.12)			
UK	0.24**	–				0.17**	0.81***	0.91	0.01	–0.12
	(0.11)					(0.08)	(0.05)			
BE			0.20***	0.14**	0.46***	0.22	0.09	0.90	0.06	–0.79
			(0.08)	(0.06)	(0.10)	(0.14)	(0.10)			
NL	0.20**				–	0.32***	0.83***	0.92	0.01	–0.09
	(0.08)					(0.10)	(0.05)			
		0.26***			–	0.32***	0.67***	0.92	0.01	–0.25
		(0.07)				(0.10)	(0.07)			
ES			0.42**	0.43***		0.65***	0.08	0.77	0.09	–0.55
			(0.16)	(0.11)		(0.20)	(0.14)			
SE	0.14*					0.30***	0.88***	0.91	0.002	–0.03
	(0.07)					(0.07)	(0.05)			
AU	0.20**					0.57***	0.70***	0.86	0.01	–0.11
	(0.08)					(0.11)	(0.09)			

The choice of foreign market to be included in the regression for each country was based on the previous results. In order to keep the analysis as simple as possible, only one foreign market was allowed to have an influence. On the basis of the previous results, the United States was included in the regressions for Italy and Canada while the United Kingdom was used in the remaining ones.[147]

The period of high volatility so defined essentially covered February to June 1994. For those countries for which the United States was used as benchmark, a few observations fell outside it. Because of the common movements in volatility, it actually made little difference which specific foreign country was chosen in the pairwise comparison.

The result are consistent with the hypothesis that spillovers are indeed larger when volatility is especially high, regardless of whether

[147] The results with Germany replacing the United Kingdom for the continental European markets were very similar.

Table AVI.7

Unconditional spillovers and the level of volatility: regression results[1]

	Foreign country		Domestic country		Dummy[2]	Tests (% probability)[3]		$\bar{R}^2$	DW
	VOL_H[4]	VOL_L[5, 6]	VOL_H[4]	VOL_L[5]		foreign H=L[7]	domestic H=L[8]		
DE	0.29***	0.01	0.51***	0.91***	1.99***	0.01***	0.00***	0.96	1.87
	(0.08)	(0.03)	(0.10)	(0.03)	(0.20)				
FR	0.63***	0.16**	0.22**	0.70***	2.67***	0.2***	0.04***	0.89	2.30
	(0.13)	(0.07)	(0.11)	(0.07)	(0.33)				
UK	0.25*	0.03	0.66***	0.91***	1.47***	13.5	7.5*	0.93	2.19
	(0.15)	(0.05)	(0.14)	(0.03)	(0.38)				
IT	0.30**	0.03	0.52***	0.82***	1.70***	5.6*	1.9**	0.89	2.04
	(0.14)	(0.06)	(0.12)	(0.05)	(0.33)				
CA	0.14	0.01	0.75***	0.87***	0.98***	29.0	27.2	0.91	1.85
	(0.14)	(0.05)	(0.11)	(0.03)	(0.34)				
BE	0.24***	0.07*	0.61***	0.85***	1.22***	1.6**	0.6***	0.92	2.25
	(0.07)	(0.04)	(0.09)	(0.04)	(0.27)				
NL	0.27***	0.06*	0.55***	0.89***	2.12***	0.1***	0.00***	0.96	1.97
	(0.06)	(0.03)	(0.08)	(0.04)	(0.25)				
ES	0.35*	0.02	0.26	0.57***	2.16***	7.2*	8.8*	0.65	2.24
	(0.21)	(0.07)	(0.20)	(0.07)	(0.59)				

[1] Estimated through the SURE technique. [2] Constant dummy equal to 1 in the high volatility period and to zero otherwise. [3] Test of the null hypothesis that the corresponding lagged volatility coefficients are the same. [4] Lagged volatility in the period of high volatility. [5] Lagged volatility in the period of lower volatility. [6] The hypothesis that all foreign lagged volatility parameters in the period of lower volatility are jointly zero is accepted (probability = 68.2%). [7] The joint hypothesis that the corresponding parameters are the same in each country is rejected (probability = 0.001%). [8] The joint hypothesis that the corresponding parameters are the same in each country is rejected (probability = 0.000%).

proximate market movements are considered in the domestic volatility process or not (Table AVI.7 and AVI.8). In all countries the spillover coefficient is considerably larger in the period of high volatility; for its part, the persistence parameter is greatly lower. In other words, spillovers are larger but fade away more quickly. Considering each country in isolation, the difference is generally statistically significant; the United Kingdom and Canada are exceptions. As applied to the *system of equations as a whole*, the two general hypotheses that spillover parameters are larger and that persistence is lower when volatility is high are each comfortably accepted at the 1% level, irrespective of whether proximate market movements are taken into account. Nor is it possible to reject the hypothesis that in the lower volatility period spillovers are non-existent.

Table AVI.8

Conditional spillovers and the level of volatility: regression results[1]

	Foreign country		Domestic country		Dummy	ΔR	Tests (% probability)		$\bar{R}^2$	DW
	VOL_H	VOL_L[2]	VOL_H	VOL_L			foreign H=L[3]	domestic H=L[4]		
DE	0.29***	0.01	0.49***	0.88***	1.74***	0.27***	0.1***	0.01***	0.96	1.88
	(0.08)	(0.03)	(0.10)	(0.03)	(0.20)	(0.05)				
FR	0.20*	–0.02	0.55***	0.81***	2.25***	0.35***	4.4**	2.6**	0.90	2.29
	(0.11)	(0.06)	(0.10)	(0.05)	(0.41)	(0.08)				
UK	0.25	0.03	0.66***	0.91***	1.52***	0.12**	14.0	7.6*	0.93	2.27
	(0.16)	(0.05)	(0.14)	(0.03)	(0.38)	(0.05)				
IT	0.23*	0.01	0.56***	0.81***	1.48***	0.28***	9.1*	3.5**	0.82	2.15
	(0.13)	(0.06)	(0.11)	(0.05)	(0.33)	(0.06)				
CA	0.16	0.01	0.74***	0.87***	1.08***	–	30.3	26.8	0.91	1.87
	(0.14)	(0.05)	(0.11)	(0.03)	(0.34)					
BE	0.24***	0.05	0.58***	0.85***	1.26***	0.24***	1.1**	0.3***	0.93	2.20
	(0.07)	(0.03)	(0.08)	(0.04)	(0.25)	(0.05)				
NL	0.26***	0.07**	0.56***	0.86***	1.86***	0.21***	0.3***	0.02***	0.97	1.96
	(0.06)	(0.03)	(0.07)	(0.03)	(0.25)	(0.05)				
ES	0.30	0.01	0.31	0.57***	1.72***	0.27***	15.7	19.4	0.67	2.31
	(0.21)	(0.07)	(0.20)	(0.07)	(0.65)	(0.09)				

[1] Estimated through the SURE technique. See Table AVI.7 for an explanation of the headings. [2] The hypothesis that all foreign lagged volatility parameters in the period of lower volatility are jointly zero is accepted (probability = 68.1%). [3] The joint hypothesis that the corresponding parameters are the same in each country is rejected (probability = 0.1%). [4] The joint hypothesis that the corresponding parameters are the same in each country is rejected (probability = 0.002%).

Annex VII – Econometric evidence on foreign (dis)investments

The basic procedure to investigate the relationship between implied bond yield volatility and foreign (dis)investments was analogous to that already employed to examine the impact of proximate market movements. Given the widespread evidence of asymmetries, net sales and net purchases of government bonds by non-residents were treated separately. These were then added to the domestic process for implied bond yield volatility, defined to include also the impact, if any, of proximate market moves. The frequency of the observations was monthly. Sales/purchases during a given month were related to volatility *at the end* of that month.[148]

The data for foreign investments approximated as closely as possible transactions in *domestic-currency fixed-rate* government bonds. Statistics for six countries were available: Germany, France, Italy, Canada, Belgium and Spain.[149] In the case of Canada certain estimates had to be made.

The regression results indicate that it is possible to discern an impact of foreign activity in some countries: Germany, France and Italy (Table AVII.1). In Canada, Belgium and Spain, by contrast, the relevant variables are not statistically significant.

In those markets where a relationship can be detected, the link survives the presence of proximate market movements. In fact, the influence of market moves becomes smaller. In the case of France, it is difficult to distinguish between the two: taken individually, the corresponding coefficients are not statistically significant; taken jointly, a clear relationship emerges. These results suggest that foreign disinvestments are an *additional* effect that works only partly through any induced change in the level of yields.

The evidence also points to an asymmetry: net sales appear to have a stronger impact that net purchases. This is true even in those two continental European markets where the relationship is not statistically significant. In the case of France, however, although the coefficient for sales is almost twice as high as that for purchases, it is not possible to reject the hypothesis that the two are the same. If so, it would be net transactions, regardless of their type, that would matter.

[148] The market movement was specified accordingly. Estimates based on month-average figures were similar.

[149] For Italy, the numbers relate to BTPs only; for France to both OATs and BTNs.

Table AVII.1

Volatility and foreign (dis)investments: regression results[1]

	Const.	Purchases	Sales	Own lag	ΔR	Tests (% probability) purchases =sales	$\bar{R}^2$	$\Delta\bar{R}^2$	DW	Sample period
DE	1.12	0.06*	0.30***	0.78***	0.40**	0.2***	0.92	0.04	2.56	9:92–3:95
	(0.82)	(0.03)	(0.08)	(0.06)	(0.18)					
	0.99	0.05	0.41***	0.83***		0.00***	0.90	0.11	2.61	
	(0.90)	(0.04)	(0.08)	(0.08)						
FR	2.09	0.05**	0.08[2]	0.77***	0.17[2]	49.4	0.78	0.00	2.70	1:93–4:95
	(1.82)	(0.02)	(0.06)	(0.13)	(0.31)					
	1.50	0.05**	0.10***	0.82***		12.1	0.79	0.11	2.77	
	(1.35)	(0.01)	(0.03)	(0.09)						
IT	6.93***	0.07**	0.37*	0.50***	0.26*	9.6*	0.66	0.03	1.57	9:92–2:95
	(1.57)	(0.03)	(0.18)	(0.11)	(0.15)					
	6.68***	0.07*	0.47**	0.52***		5.4*	0.62	0.10	1.69	
	(1.86)	(0.04)	(0.21)	(0.13)						
CA	4.13	–0.17	–0.24	0.77***	–	83.8	0.62	–0.02	1.16[1]	9:92–4:95
	(1.99)	(0.33)	(0.19)	(0.11)						
BE	2.20**	9.7	19.0	0.79***	–	55.1	0.78	–0.01	1.92	9:92–12:94
	(0.99)	(10.2)	(12.6)	(0.09)						
ES	6.45**	–0.27	1.38	0.57***	0.35**	41.3	0.46	–0.02	1.68	12:92–3:95
	(2.60)	(0.58)	(2.06)	(0.15)	(0.13)					

[1] Adding an additional lagged value for implied volatility eliminates the residual serial correlation but does not alter the basic results with respect to foreign (dis)investments. [2] Foreign sales and declines in bond yields are jointly significant at the 5% level.

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