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Abstract

This paper analyses the impact effects of changes in central bank rates on the term structure of interest rates in nine industrial countries over a period extending from 1987 to 1995. We try to identify an immediate channel linking central bank actions to inflation expectations. Our main results are: i) on average changes in central bank rates had a substantial impact not only on short-term but also on long-term segments; ii) while the responses of short-term rates were equivalent across countries, the reactions of long-term rates differed markedly; three groups of European countries can be identified, according to whether, over the period considered, central bank rates rises induced increases (Italy, Spain, Sweden, United Kingdom), no change (France) or decreases (Germany, the Netherlands and Belgium) in long-term forward rates; iii) such differences are closely correlated with past inflation records and, to a lesser extent, with public finance imbalances; this outcome may reflect different degrees of credibility of the long-term anti-inflationary commitment of monetary policy, resulting from the market perception not only of the central bank's intentions, but also of the long-run sustainability of the monetary policy stance; iv) the estimated relationships show significant changes over time in the US; signs of modification are perceptible also in Italy in 1995. In the recent experience of both countries increases in central bank rates have brought about decreases in long-term forward rates.

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Contents

1. Introduction and main conclusions[1]

Economic theory suggests that monetary policy actions can affect the term structure of interest rates through the liquidity effect, which should only impinge on the short-term segments, and by influencing inflation expectations and risk premia, which in turn affect the long-term segments.

In this paper we analyze the *impact* effect of changes in central bank rates on the term structure of interest rates in nine main industrial countries over a period extending from 1987 to 1995. We look at the reactions of market interest rates over the weeks immediately surrounding changes in monetary policy rates. In particular, we try to identify a direct channel linking central bank actions to movements in inflation expectations. We do not address other effects of monetary policy on inflation and inflation expectations, which unfold gradually over the longer run.

The analysis revolves around the following issues:

i) Do changes in monetary policy rates have an impact on long-term segments of the term structure of interest rates in addition to that on the short-term end?
ii) Do such responses - which presumably reflect mainly the reactions of inflation expectations - differ across countries and over time?

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iii) How can these differences be explained?

The key conclusions of the paper can be summarized as follows.

i) *Over the periods considered*, changes in the central banks' rates displayed *on average* a substantial impact not only on the short-term but also on the long-term segments of the term structure of interest rates.

ii) While the responses of short-term rates were equivalent across countries, the reactions of long-term rates differed markedly. In particular, three groups of European countries can be identified, according to whether central bank rates rises induced increases (Italy, Spain, Sweden, United Kingdom), no change (France) or decreases (Germany, the Netherlands and Belgium) in long-term *forward* rates.

iii) *Cross-country differences* in the impact of monetary changes on long-term rates are highly correlated with *past inflation records*. They also display a significant correlation with potential macroeconomic constraints on monetary policy, especially those arising from public finance imbalances. This piece of evidence suggests that different reactions of inflation expectations may reflect different degrees of *credibility* of the long-term anti-inflationary commitment of monetary policy. This, in turn, results from the market perception not only of the central bank's intentions, but also of the long-run sustainability of the monetary policy stance, related to other components of economic policy, especially fiscal policy.

iv) In the United States, the estimated response of market rates to monetary actions shows significant changes over time. In particular, in the last two years rises in the Federal funds rate target have resulted in reductions of long-term forward rates, in contrast with the increases registered, on average, in previous occasions. Signs of a similar modification emerged in Italy in 1995, where a sharp tightening of monetary policy began in 1994. These changes over time may have been originated by shifts in the market perception of the determination of central banks to fight inflation and of the long-run sustainability of such intentions.

The paper is organized as follows: in Section 2 we provide a simple framework for understanding how monetary policy moves may affect the term structure of interest rates. In Section 3 we discuss possible interpretations of long-term rates responses to central banks' actions, focusing on the anti-inflationary commitment of monetary policy. In Section 4 we first estimate average responses of market rates to policy rate changes over the sample periods and then we analyze possible determinants of cross-country differences in these responses. Finally, we check whether these responses change through time in Germany, Italy and the United States.

2. Monetary policy and the term structure of interest rates

The forward rate at time t with settlement at time t' and maturity at T, $f(t,t',T)$, can be written as the sum of three components: i) an "expected real rate component"; ii) an "expected inflation component"; iii) a "risk premium component":

(1) $f(t,t',T) \equiv E_t r(t',T) + E_t \pi(t',T) + \xi(t,t',T)$,

where the first component, $E_t r(t',T)$, is the expected real rate of interest between t' and T; the second, $E_t \pi(t',T)$, is the expected rate of inflation between t' and T; the third, $\xi(t,t',T)$, is the risk premium.[2] For simplicity, we assume that the risk premium can be linearised into the sum of two components: an inflation risk premium $\vartheta(t,t',T)$ and a non-inflation risk premium $\eta(t,t',T)$. The distinction between expected inflation and inflation risk premium is not actually crucial for our purposes, since, as Svensson (1994) points out, an increase in either one is undesirable for a central bank whose goal is price stability. We will thus refer to the sum of the two effects as to the "*inflation component*" of forward rates, using the notation $E_t^* \pi(t',T)$. Expression (1) can thus be rewritten as:

(2) $$f(t,t',T) \equiv E_t r(t',T) + E_t^* \pi(t',T) + \eta(t,t',T),$$

where $$E_t^* \pi(t',T) \equiv E_t \pi(t',T) + \vartheta(t,t',T).$$

As the settlement date approaches (t→t') and the maturity shortens (T→t), the forward rate tends to the instantaneous short-term spot rate, f(t)→i(t). Over such a short horizon, the inflation rate is sticky at its current level π, while the real interest rate adjusts fully to that implied by the level of the short-term rate chosen by the central bank:

[2] For precision, a convexity term should be included, but we omit it as not relevant for our analysis.

(3) $i(t) \cong r(t) + \bar{\pi}$.

Moving along the forward curve, as t′ increases towards longer horizons, the expected real rate tends to its long-term, structural value, $\bar{\bar{r}}$, while inflation stickiness diminishes and inflation expectations tend to adjust to the level implied by the present and expected monetary policy stance:

(4) $f(t,t',T) \approx \bar{\bar{r}} + E_t^* \pi(t',T) + \eta(t,t',T)$.

Positing that the *non-inflation* component of the risk premium is not systematically affected by monetary policy moves, which we believe to be a realistic assumption, one may interpret forward rate *changes* in response to monetary policy rate movements as variations of the inflation component, the expected real rate, or a combination of the two. In formal terms, this implies that the mean of the first difference of the forward rate (2), conditional on the change of the monetary policy rate i^P, is given by:

(5) $E_t[\Delta f(t,t',T) | \Delta i^P(t)] = E_t[(E_t \Delta r(t',T) + E_t^* \Delta \pi(t',T)) | \Delta i^P(t)]$,

since we assume the conditional mean of the change in the non-inflation risk premium to be nil: $E_t[\eta(t,t',T) | \Delta i^P(t)] = 0$. This allows us to estimate the regression:

(6) $\Delta f(t,t',T) = \alpha + \beta \Delta i^P(t) + \varepsilon(t)$,

assuming that changes in the non-inflation risk premium are captured by the error term $\varepsilon(t)$. The interpretation of the coefficient β in this regression, which is very similar to the regression models that will be estimated in Section 4, varies depending on which segment of the forward curve is considered. In the case of the *short-term spot rate* (equation (3)), this coefficient would capture the so-called *liquidity effect*. As prices are constant over a very short horizon, invariant to the central bank move, the entire change in the instantaneous rate is equivalent to a change in the instantaneous *real* rate of interest:

(7) $\Delta i(t) \cong \Delta r(t) = \alpha + \beta \Delta i^{P}(t) + \varepsilon_{t}$.

Moving along the forward curve towards medium-term horizons, the coefficient β would instead represent a combination of the effects of monetary policy actions on both the real rate and the inflation components. Further on, in the regression of long-term forward rates, it would measure the inflation effect alone:

(8) $\Delta f(t,t',T) \cong \Delta E_{t}^{*}\pi(t,t',T) = \alpha + \beta \Delta i^{P}(t) + \varepsilon_{t}$.

In Section 4 we will split the forward rate curve into three segments: the "near-term", up to one year, whose changes are likely to reflect mainly the liquidity effect; the "intermediate-term", one to five years, which should reflect expected changes both in the real rate and the inflation component; and the "long-term", five to ten years, which should provide clearer indications of the impact of monetary actions on (long-term) inflation expectations.

Yields to maturity - which convey the information available in the term structure differently from forward rates - are also considered. Yields on par bonds (like the swap rates considered in our empirical analysis) paying (k) coupons (c) can be approximated by a weighted average of the current spot rate and the implicit forward rates, with the largest weight on the current rate and decreasing weights from near-term to long-term forward rates:[3]

(9) $$i(k,t,T) \cong \sum_{j=0}^{k} v(k,j) f(t,t+j,t+j+1) ,$$

with $$v(k,j) = \frac{\sum_{i=j}^{k} c\frac{1}{i} + \frac{1}{k}}{(1+kc)} ,$$

$$i(t,t+1) = f(t,t+j,t+j+1) \quad \text{for} \quad j = 0 ,$$

and $$\sum_{j=1}^{k} v(k,j) = 1 .$$

Thus the reaction of yields to maturity combines the liquidity effect on the short-term spot rate and the inflation effect on forward rates, making it hard to discriminate between the two. In particular, the response of long-term yields to maturity does not allow one to infer whether inflation expectations have risen or fallen, as yields are likely to move in the same direction as the central bank rate due to liquidity effects. Nevertheless, we also performed an analysis of yields reactions to monetary shocks, for two reasons. First, it permits comparison of our results with those of most previous empirical works, which were based on these rates, and second it provides useful information for international comparisons: if the liquidity effects are of the same magnitude across countries, differences in long-term yields responses could still be

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3 For this, see Radecki and Reinhart (1994).

interpreted as the result of different effects on the average inflation component over the maturity they refer to.

3. Some guidelines for interpretation

As was anticipated in the previous section, the empirical analysis splits the term structure into a "near", an "intermediate" and a "long" term segment. The liquidity effect should dominate in the first segment and we do not expect it to differ significantly between countries. In the intermediate segment, two effects should combine: the effect on expectations of future real short-term rates and the inflation effect, on inflation expectations and inflation risk premium. In the long-term end the latter effect should prevail. Although the distinction among these effects is crucial, it has not always received sufficient attention in the literature. Radecki and Reinhart (1994) and Roley and Sellon (1995), for instance, do not fully recognize that any given effect of a policy move on nominal rates may result from various possible combinations of changes in real rate and inflation expectations. As we shall see at the end of this section, this may generate interpretations of long-term rate reactions that differ from our own. Our analysis concentrates on the long-term segment of the term-structure, as we are interested mainly in the effects of policy moves on inflation expectations, which we believe to be dominant in this segment and to hinge on *monetary policy credibility*, a concept that does not coincide with that of *central bank credibility*, as it depends not only on the central bank's resolution to fight inflation but also on constraints to its actions, mainly arising from fiscal policy and the structure of the economy.

This interpretation is consistent with other analyses, such as those of Goodfriend (1993), King (1995) and Volcker and Gyohten (1992), who emphasize that the impact of monetary policy on expected inflation depends on the market perception of the long-term commitment of monetary policy to low inflation. Volcker and Gyohten, for instance, explain the increase in long-term interest rates which followed the discount rate rises in 1980-81 (a turning point in the Fed's policy) in the following terms: "One telltale sign of our difficulties was that when short-term interest rates rose, as anticipated, long-term rates did, too. If the markets were convinced that inflation would be coming down, presumably that would not have happened, at least not for long" (p. 170). Also Goodfriend's more optimistic assessment of the effects of the tightening in the first part of 1980 focuses on monetary policy credibility. He notices that, with the "enormous 3 percentage point increase of the monthly average funds rate in March ... the long rate hardly moved in response, suggesting that the positive effect on the long rate of the aggressive tightening was offset by a decline in expected inflation. Moreover, the long rate actually came down by 0.9 percentage points in April even as the Fed pushed the funds rate up another 0.4 percentage points, suggesting that the Fed had already begun to win credibility for its disinflationary policy" (p. 11).

We first consider the case of a rise in the central bank rate, and suggest a possible interpretation of different reactions of long-term interest rates in terms of monetary policy credibility. We then consider the case of a reduction in the central bank rate, suggesting that considering the reaction of long-term rates to both rises and cuts at once may help us to distinguish three cases - a "super-credible", a "credible" and a "non-credible" monetary policy - and may also indicate possible alternative reasons for the differing credibility of monetary policy between countries.

Consider first the case of a central bank rate rise.[4] If the policy move is seen by the market as effective in lowering the underlying, long-term inflation rate, it should determine a reduction of long-term inflation expectations and thus of long-term forward rates. The opposite would occur if the market perceived the policy move as unsustainable and thus doomed to eventual reversal. A well known case is the "unpleasant monetarist arithmetic" of Sargent and Wallace (1981), referring to a monetary restriction that makes the growth of public debt unsustainable in the presence of persistent primary deficits, eventually inducing the government to force even an "independent" central bank to be more inflationary in the future to repay the debt. Constraints could also stem from other macroeconomic factors, such as high unemployment, that monetary tightening may help drive to unsustainable levels.[5] In the presence of such constraints, an increase in official rates could produce an adverse effect, determining a rise in long-term inflation expectations and thus in forward rates. Such an adverse result could also occur in the absence of a "Sargent-Wallace effect", as monetary policy could lack credibility owing to other factors, such as the lack of independence of the central bank or a market perception that monetary authorities are insufficiently committed to price stability. These factors are particularly relevant when the markets view a policy move,

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4 We assume that a policy rate rise is always perceived by the market as a policy tightening and vice-versa for a cut. This assumption, although perhaps not appropriate for each single case (for instance when the rise is smaller than expected by the market), seems to be reasonable "on average" and is adopted also in the empirical section of the paper.

5 In this regard, see Drazen and Masson (1994).

say a rate rise, as revealing the central bank's information on the inflation risks in the economy, in particular when the rise represents a turning point in the monetary stance. In this case the effect on inflation expectations will depend crucially on confidence in the central bank's ability to keep the monetary policy stance tight enough, and for long enough, to counter these risks.

Considering the reaction of long-term rates to central bank rate cuts[6] together with the effects of the increases enables us to distinguish four possible cases:

i) long-term rates decrease or remain unchanged in the case of a central bank rate rise and decrease or remain unchanged in the case of a cut. In the estimates of regressions like that of equation (6), which we perform in Section 4, this would imply a negative or nil coefficient for rises and a positive or nil coefficient for cuts. We call this the case of "super-credible" monetary policy, whose actions *never* induce a rise in long-term inflation expectations.

ii) Long-term rates decrease in the case of a rise but increase in the case of a cut, which implies that the coefficient in the regression would be symmetric, negative in both cases. We call this as simply "credible" monetary policy, whose moves determine a "textbook" relationship with inflation expectations, which fall when policy is tightened and rise when policy is eased.

[6] See footnote 4.

iii) Long-term rates increase in the case of a rise and decrease in the case of a cut, which implies that the coefficient in the regression would be symmetric, positive in both cases. This case suggests that monetary policy is not credible due to some "Sargent-Wallace effect", i.e. that inflation expectations rise when monetary policy is tightened and viceversa when policy is eased.

iv) Long-term rates increase both in the case of a rise and in the case of a reduction, which implies that the coefficient in the regression would be asymmetric, positive for rises and negative for cuts. This case suggests that monetary policy is not credible for reasons other than "Sargent-Wallace effects", such as a market perception of insufficient central bank commitment to price stability.

A relevant implication of this scheme for the empirical analysis is the need to investigate the presence of asymmetries by running separate regressions for central bank rate rises and cuts. Only in the absence of such asymmetries could results for a unique regression including both rises and cuts be considered significant.

Interestingly, the foregoing interpretation provides viewpoints which may differ significantly from those of other authors. According to Roley and Sellon (1995), for instance, increases in the central bank rate may result in reductions of long-term rates "if investors believe a current policy action will be fully offset and ultimately reversed in the future" (p. 80), while a positive relationship would be observed were the policy rate rise expected to persist. In our interpretation, the opposite may hold. In the case of a rate rise, for instance, a negative relationship may well emerge precisely because the markets do believe that the anti-inflationary stance will be maintained, which reduces inflation expectations (cases i and ii). Conversely, expectations of a future reversal of the policy would determine a positive relationship (cases iii or iv). The contrast between the two interpretations, however, is less substantial than it might appear, as both the real rate component (which is what Radecki and Reinhart and Roley and Sellon implicitly focus on when they interpret market rate reactions in terms of persistence of monetary policy) and the inflation component (which is what we concentrate on when looking at the long-term segments) combine to determine expectations about future nominal rates, hence the reaction of present rates. What is crucial in our analysis is that the two components may have different weights in different segments of the term structure, and thus that market rate reactions to policy moves may take different explanations depending on the time horizon.

4. Empirical analysis

In this section we examine the impact of changes in monetary policy rates on the term structure in nine countries: Belgium, France, Germany, Italy, the Netherlands, Spain, Sweden, the United Kingdom and the United States.

4.1 *Data description*

Monetary policy innovations can be measured by a number of variables. Shiller, Campbell and Schoenholtz (1983) look at market rate reactions to money growth announcements in the United States for 1979-1982, when the Federal Reserve was targeting non-borrowed reserves and quantity variables played a dominant role in monetary policy. More recent papers consider the effects of changes in key monetary policy rates set by the central banks, namely the Federal funds target in the US (Cook and Hahn, 1989; Radecki and Reinhart, 1994; Roley and Sellon, 1995) and the base rate in the United Kingdom (Dale, 1993). We adopted the latter approach, as changes in central bank rates provide a more appropriate measure of variations of the monetary policy stance in the countries and over the periods considered.

For some countries we considered changes in two monetary policy interest rates. In fact, a number of central banks have two key policy rates: i) "official rates" such as the discount rate or the Lombard rate, which are moved less frequently and provide a strong signalling content on the medium-term stance of monetary policy; ii) rates that are changed more frequently - as they concern operations aiming at regulating liquidity, such as repos - whose variations do not necessarily signal the medium-term orientation. Changes in official rates are likely to transmit the monetary impulse through their impact on expectations, thus mainly affecting the long-term segments of the term structure. The effect of other policy rates should be limited to the short-term and medium-term segments.

Our choice of the monetary policy rates was consistent with a number of official and research papers drafted by central bank officers and by international organizations. We used one policy rate for Spain (auction rate), Sweden (marginal lending rate), the United Kingdom (base rate) the US (Federal funds rate target) and the Netherlands (special loans rate). Two policy rates were considered for Belgium (discount rate and central rate), France (intervention rate and overnight rate),[7] Germany (average of discount rate and Lombard rate; repo rate) and Italy (discount rate and repo rate).

We analyze the effect of changes in monetary policy rates both on yields to maturity and on forward rates. Yields to maturity refer to the 1-year Libor rate (1-year domestic interbank rate for Spain and Sweden), and 3, 5 and 10-year euro-market swap rates. All swap rates correspond to yields to maturity of par-bonds paying a semi-annual coupon. We also analyze the behaviour of forward rates implicit in euro-market rates:[8] an intermediate-term forward rate, the 1-year-ahead 4-year forward rate (1-5 year in the tables) and a long-term forward rate, the 5-year-ahead 5-year forward rate (5-10 year). All data refer to end-of-week observations. We analyze the reaction of market rates in the weeks surrounding policy rate moves. This time interval, larger than in most of the previous literature[9] - with the exception of Roley and Sellon (1995) who use a larger interval[10] - aims at capturing the *total* impact response of interest rates, including *both anticipatory and learning effects*.

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[7] We used the overnight rate - which is closely managed by the Bank of France - as it captures changes in monetary policy conditions especially in periods in which the *pension* facility is temporarily suspended by the Bank of France.

[8] The methodology follows Drudi and Giraldi (1991) and is described in the Appendix.

[9] Shiller et al. (1983), Cook and Hahn (1989), Radecki and Reinhart (1994) and Dale (1993) consider market rate responses over a very narrow interval, either the same day as the monetary action or at most the days immediately surrounding the event.

[10] The average coefficients we estimate for the whole sample period for the US are consistent with theirs.

By looking at the responses of market rates to monetary policy rates we can identify, in particular, an immediate channel linking central bank actions to inflation expectations. This methodology, instead, cannot measure the full effect of monetary policy on interest rates, as it neglects its influence on inflation and inflation expectations, which is exerted with longer lags and does not stem immediately from the policy move.

The periods examined vary across countries, depending on data availability for each country. A longer period, from 1987 to 1995, is considered for Germany, the United Kingdom and the United States; a shorter one, from 1991 to 1995, for Belgium, France, Italy, the Netherlands and Spain; and from 1992 to 1995 for Sweden.[11] Thus, for most countries, the data refer to a relatively short period.[12]

Figures 1a-1c show the evolution of the monetary policy rates used in the empirical analysis, of the 1-year rate and of the 5-year-ahead 5-year forward rates over the sample selected for the nine countries.

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[11] Belgium: June 1991-December 1995. France: January 1991-December 1995. Germany: April 1987-December 1995. Italy: March 1991-December 1995. Netherlands: June 1991-December 1995. Spain: January 1991-December 1995. Sweden: November 1992-December 1995. United Kingdom: April 1987-December 1995. United States: April 1987-December 1995.

[12] The data sources are the following:
- For Italian data: Bank of Italy data base.
- For monetary policy rates for all other countries: BIS data base.
- For euro-market rates: Data-Stream for Germany, Netherlands, UK and USA; Bloomberg for Belgium, France, Spain and Sweden.
- The Federal funds targets are from Rudebusch (1995).

4.2 *Average market rate responses to monetary policy moves*

We ran two sets of regressions, of the type of equation (6). For each country, we regressed first differences of the selected market rates on: i) their lagged values; ii) contemporaneous, lead and lagged values of changes in monetary policy rates (as noted, using one or two policy rates depending on country); iii) contemporaneous and lagged values of changes in the US rate with the same maturity as the endogenous rate, to capture interest rate movements related to international shocks rather than to changes in domestic policy rates:[13]

$$\Delta i_t = \alpha_0 + \sum_{i=-1}^{n} \beta_i \Delta i_{t-i}^{P1} + \sum_{j=0}^{n} \delta_j \Delta i_{t-j}^{*} + \sum_{k=1}^{n} \varphi_k \Delta i_{t-k} + \varepsilon_t , \quad (10)$$

or

$$\Delta i_t = \alpha_0 + \sum_{i=-1}^{n} \beta_i \Delta i_{t-i}^{P1} + \sum_{i=-1}^{n} \gamma_i \Delta i_{t-i}^{P2} + \sum_{j=0}^{n} \delta_j i_{t-j}^{*} + \sum_{k=1}^{n} \varphi_k \Delta i_{t-k} + \varepsilon_t , \quad (10')$$

depending on whether one or two policy rates were included in the regression, with:

$i_t \equiv$ endogenous market rate; $i_t^{P1} \equiv$ official rate; $i_t^{P2} \equiv$ other central bank rate;

$i_t^{*} \equiv$ US rate; $\varepsilon_t \equiv$ error term.

One lead of the policy rate was included in all regressions, to capture not only market anticipation of the policy move, but also possible contemporaneous reactions that would otherwise not be captured by the data (for instance, when a policy move is decided and announced on Friday afternoon, with the markets still open, but

[13] No "foreign" rate was included in the regression for the US.

officially recorded on the following Monday).[14] Zero to two lags were included for all variables; the number of lags, which was chosen for each regression according to the Schwarz criterion and the absence of residuals serial correlation, is indicated for each country in the footnotes to Tables 2a-2i.

We also tested for the presence of possible asymmetries in the effects of policy rate rises and cuts, by estimating a regression including rises and cuts separately:

$$(11) \qquad \Delta i_t = \alpha_0 + \sum_{i=-1}^{n} \beta_i^{\uparrow} \Delta i_{t-i}^{P1\uparrow} + \sum_{i=-1}^{n} \beta_i^{\downarrow} \Delta i_{t-i}^{P1\downarrow} + \sum_{j=0}^{n} \delta_j i_{t-j}^{*} + \sum_{k=1}^{n} \varphi_k \Delta i_{t-k} + \varepsilon_t \, .$$

The main results are summarized in Table 1, which reports for each country the sum of the coefficients of contemporaneous, lead and lagged values of policy rate changes in regression (10) for the Netherlands, Spain, Sweden, and the US and of regression (10') for Belgium, France, Germany and Italy. Coefficients relative to monetary policy rates rises, according to model (11), are presented for the UK and France, as for these two countries there is evidence of asymmetric effects of cuts and rises on the 5-10 year forward rate (although statistically significant only for the UK). For other countries we found no significant difference between the coefficients of rises and those of cuts and we reported the coefficient of the more general model (10-10'). Detailed results for each country are shown in Tables 2a-2i, one for each country.

The main results may be summarized as follows:

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[14] The results we obtain by omitting the lead of the policy rate are not qualitatively different from those presented in the paper.

i) In the regression for the 1-year Libor rate, the coefficient is not significantly different from unity anywhere, indicating that over the sample periods considered the effect of monetary policy actions on short-term rates was similar in all countries.

ii) The impact effect of changes in central bank rates was not limited to short-term rates, as it extended through longer segments of the term structure.

iii) The response of medium and long-term yields to maturity was both substantial and significant in all the countries and for all the maturities, even the longest. The estimated coefficients are generally larger than those of most previous empirical works. For instance, in Dale (1993) for the UK and in Radecki and Reinhart (1994) for the US, changes in the monetary policy rate do not display a systematic influence on interest rates with maturities longer than 5 years. Our results for the US are closer to those obtained more recently by Roley and Sellon (1995). The difference with previous results may be explained, as the latter suggest, by the larger time interval used both in their paper and in ours to measure the effects of policy rate changes (see footnotes 9 and 10), which captures a larger portion of market rate responses.

iv) Estimated coefficients for the long-term segments - which reflect mainly the inflation effect - display significant differences across countries. In particular, for the 5-10 year forward rate, whose response provides the sharper signal of the impact of monetary policy moves on inflation expectations (see Section 2), three groups of European countries can be identified.

- In one group, comprising Germany, the Netherlands and Belgium, the coefficient is negative (ranging from -0.14 to -0.19), implying that in the period considered the average impact of monetary policy rate rises was a decrease of long-term forward rates. The above-mentioned symmetry between the effects of rises and cuts in policy rates suggests for these countries a "textbook" effect of policy rate changes on inflation expectations (see Section 3), which tend to decrease when monetary policy is tightened and to increase when it is eased. Interestingly, there is no "super-credible" monetary policy; that is, no central bank whose moves never increase inflation expectations.
- In a second group, comprising Italy, Spain, Sweden and the UK, central bank rate rises determined significant increases in long-term forward rates (the coefficients are positive, ranging from 0.20 to 0.24). Results are symmetric for rises and cuts in the first three countries, suggesting that, during the sample period considered, a "Sargent-Wallace effect" may have influenced the responses of market rates. As mentioned, results are asymmetric for the UK, as is shown by Table 2g: the sign for cuts is negative (although not statistically significant), showing that long-term forward rates tended to increase also when the base rate was lowered (the difference between the coefficients for rises and cuts is statistically significant; Wald test in the table). This may signal a lack of confidence in the anti-inflationary commitment of monetary policy rather than the presence of a "Sargent -Wallace effect".

- In France,[15] increases in monetary policy rates determined no significant changes in long-term forward rates. As in the UK, though, a negative coefficient for rate cuts (although not statistically significant) signals that on these occasions, on average, expected inflation tended to rise.

v) In countries where two policy rates are included in the regressions, they display different effects on the various segments of the term structure, especially in Germany and in Italy where the distinction between them is clear-cut. In both countries the effects of repo rates are prevalent on short-term rates, but negligible on long-term forward rates; the latter are significantly affected by "official" rates only.

vi) In the European countries considered the monetary policy rate coefficients do not display significant breaks over the sample periods. On the contrary, significant changes are found in the US. As shown in Table 1, which reports estimates for the whole 1987-1995 sample period and for the 1991-95 sub-sample, the coefficient declines significantly in the more recent period. More evidence in this regard is provided in Section 4.4.

4.3 *Interpretation of international differences*

In this section we try to identify possible determinants of international differences in the responses of long-term rates to policy rate changes. As mentioned in Section 3, our hypothesis is that different reactions over this segment may

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15 The extraordinary 10-point increase of the overnight rate in September 1992 and the subsequent fall in October were omitted from the regressions.

reflect differing credibility of the monetary policy anti-inflationary commitment, which may in turn depend both on past inflation performance and on macroeconomic constraints on the central bank's action. We thus measure the correlation of the responses first with the past inflation experience of each country and then with variables capturing possible constraints on monetary policy (fiscal imbalances, level of indebtedness, unemployment).

Ideally, we would like to run a cross-country multivariate analysis including all candidate explanatory variables; this is prevented, however, by the limited degrees of freedom available. A bivariate analysis, looking at correlations between the estimated coefficients and each explanatory variable, may be considered as a first step in the study.[16]

Past inflation performance is a candidate variable to explain monetary policy credibility, on the assumption that markets "have a long memory" and, thus, that they form their expectations, at least to some extent, by extrapolating past performance.[17] The scatter plot of the estimated effects of policy rate changes on 5-10-year forward rates (corresponding to those reported in Table 1) against past inflation is shown in the top-left panel of Figure 2a. The interpolating line is positive and the slope coefficient is significantly different from zero (t-ratio=10.91); and the explanatory value of past inflation, as summarized by the R^2 coefficient, is very high (0.95). This result seems to

[16] We did not include the estimated coefficients of the US in the analysis of this paragraph since, as mentioned, they proved to be highly unstable over the sample period.

[17] King (1995) and Ganley and Noblet (1995) use past inflation as an explanatory variable of cross-country differences in the behaviour of long-term interest rates in the 1993-94 period.

suggest that a "good" inflation record may have contributed, over the sample period, to a favourable impact of monetary policy rate rises on inflation expectations and the risk premium in Germany, Netherlands and Belgium; conversely for the other countries.

This conclusion is strengthened by a comparison of the results for the 5-10 year forward rate with those obtained for the other market rates considered (Figure 2b). The correlation with past inflation is nil for the estimated effects of policy rate changes on yields to maturity up to 3 years, which are more affected by the liquidity effect (see equation (9)). It increases slightly on 5-year swap rates and on intermediate-term (1-5 year) forward rates which, as pointed out in Section 2, may reflect expected movements both of inflation and of real rates. The correlation is high and significant for the responses of 10-year yields to maturity (t-ratio=7.36; R^2=0.90). This seems to confirm that the reactions of the short and the medium-term segments of the term structure mainly reflect liquidity effects and real interest rate expectations, while the inflation effect, which depends on monetary policy credibility and thus on past inflation performance, becomes a dominant factor over the longer-term segment of the curve.

Among macroeconomic constraints, the presence of structural public finance imbalances has been identified in the literature as a possible source of unsustainability of tight monetary policy (the "unpleasant arithmetic" of Sargent and Wallace; see Section 3); it is also often cited in the comments of financial market analysts and participants as a major determinant of long-term interest rates in some of the countries considered. To capture this effect, we considered the stock of public debt and the past record of both total and primary budget deficits as possible explanatory variables of the reactions of long-term forward rates. Among these three variables, only the primary deficit displays a significant correlation with the impact of monetary policy (Figure 2a). A possible interpretation is that differences in structural programs of fiscal consolidation are considered by financial markets more important than differences in the outstanding stock of government debt, as a continued stream of primary surpluses may signal the ability of a country to sustain the burden of a high debt in the long run. This may be the case of Belgium, which is characterized by textbook reactions of long-term rates in response to monetary shifts, notwithstanding the high public indebtedness.

We also considered the possible effect of the presence of a high short-term and variable rate total debt (both private and public) and historically high unemployment rates. In many occasions, these two variables have been considered by some commentators as representing a problem for the sustainability of tight monetary policies in some of the countries considered, particularly in the UK (short duration of the private debt), in Italy (short duration of public debt) and in France (high unemployment). A positive correlation between these factors and the effects of monetary policy seems to emerge from the scatter plots in Figure 2a, although it is statistically significant only for the short-term debt.

4.4 *Analysis of regime shifts within individual countries*

In Section 4.2 we presented estimates of the average responses of long-term interest rates to policy rate changes in each country over the sample period. The way the markets perceive the anti-inflation determination of monetary policy makers, however, may change over time, being affected by shifts in the monetary regime, possibly determined by changes in the institutional set-up of the central bank, in the strategy of monetary policy, in the exchange rate regime. Standard stability tests based on Recursive Least Squares (RLS) pointed out significant changes of the estimated relationship only in the US. In this section, for a sub-sample of countries, we also present estimates of time-varying coefficients of the contemporaneous effects of changes of the official rate on 5-10 year forward rates,[18] adopting a methodology based on the Kalman filter. Using this technique we try to detect any change in the relationship not captured by more traditional procedures. The analysis was carried forward for two of the countries for which a longer sample period was available (Germany and the US), and for Italy. The results for Italy should be interpreted with caution, owing to the relatively short sample period.

Assuming that the parameters of regression (10) evolve over time as random walks, we reformulate the model as follows:

$$(12) \qquad \Delta i_t = \alpha_{0,t} + \beta_{0,t}\Delta i_t^{P1} + \sum_{j=0}^{n} \delta_{j,t}\Delta i_{t-j}^{*} + \sum_{k=1}^{n}\varphi_{k,t}\Delta i_{t-k} + \varepsilon_t, \qquad \varepsilon_t \approx N\left(0,\sigma^2\right),$$

$$\lambda_t = \lambda_{t-1} + \xi_t, \qquad \xi_t \approx N\left(0,\sigma^2\mathbf{Q}\right),$$

[18] The coefficient of contemporaneous changes of the Federal funds rate in the US, of the discount rate in Italy and of the average between the discount rate and of the Lombard rate in Germany accounts for the largest part of the total coefficient of monetary policy rate changes shown in the tables.

where $\lambda_t \equiv (\alpha_{0,t} | \beta_{0,t} | \delta_{0,t} | \delta_{1,t} | \ldots | \delta_{n,t} | \varphi_{1,t} | \varphi_{2,t} | \ldots | \varphi_{n,t})$ and Q is a matrix representing the signal-to-noise ratio, i.e. the extent to which prediction errors made at time t are attributed to noise or, instead, interpreted as a signal that the vector λ_{t-1} must be revised. To estimate Q, it would be necessary to determine the $(2n+3)*(2n+4)/2$ parameters. Because of the limited number of degrees of freedom, Q is set exogenously at $Q=\chi I_k$, where χ is a given parameter. If χ is low, prediction errors can interpreted as noise around a relatively stable coefficient. When χ is nil, the estimating procedure coincides with RLS. On the contrary, if χ is high, large prediction errors cause a high variability of the parameters λ. The model is estimated for three values of χ: 0, 0.01 and 0.1. As mentioned, the first estimate is equivalent to that obtainable using RLS and corresponds to the polar assumption that prediction errors are entirely attributed to noise, while the parameters are fixed. In the case of the larger values of χ, 0.01 and 0.1, each central bank rate move can lead to changes in the market perception of the credibility of monetary policy, and therefore in the parameter $\beta_{0,t}$. The estimated coefficients $\beta_{0,t}$ for the three values of χ are plotted in the upper part of Figures 3a-3c. The evolution of official and 5-10 year forward rates in the corresponding periods is shown in the lower part of the figures.

In the United States (Figure 3a), the RLS coefficient diminishes progressively over time, signalling a number of structural breaks. This finding confirms the evidence presented in Table 1, which reports estimates both for the whole period 1987-1995 and for the sub-period 1991-95. The estimate with $\chi=0.1$ shows that the coefficient turns negative in 1994, when rises in the Fed funds target brought about decreases in long-term forward rates.[19] The sharp tightening in 1994 and in 1995 by the Federal Reserve, which acted preemptively in the face of emerging inflationary pressures, signalling a strong determination to fight inflation, is likely to have improved its long-term anti-inflationary credibility.[20]

In Italy (Figure 3b) the RLS estimate does not indicate the presence of structural breaks. However, some signs of modification do emerge from the estimate with values of χ greater than 0, although only very recently. The coefficient is positive in 1992 and 1993, since during the EMS crisis the impact effect of official rate rises on long-term rates was positive, while that of the subsequent cuts was negative. The coefficient was also positive during the first part of the recent tightening, which began in mid-1994,[21] as the rise in official rates in August 1994 determined a contemporaneous increase in long-term interest rates (similarly to what happened in the US in February 1994; see footnote 19). Signs of an improvement, however, emerged in

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19 Actually, unlike subsequent moves, the first increase in the Fed fund rate target in February 1994 resulted in higher long-term forward rates, as reflected also in the behaviour of the coefficient. It came after a phase of policy rate cuts (between 1989 and the second half of 1992) and a subsequent phase of stable rates. It was perceived by the market as providing new information about impending changes in official rates and about the fundamentals of the economy and induced a substantial rise in the whole term structure (the same occurred in other countries in case of inversions of the monetary stance; for instance, the reduction in the UK in the second half of 1990 and the increase in 1994; the increase in Italy in August 1994; the first increase in Spain in 1995).

20 On the evolution of the Fed's orientation, see Goodfriend (1993), Volcker and Gyohten (1992).

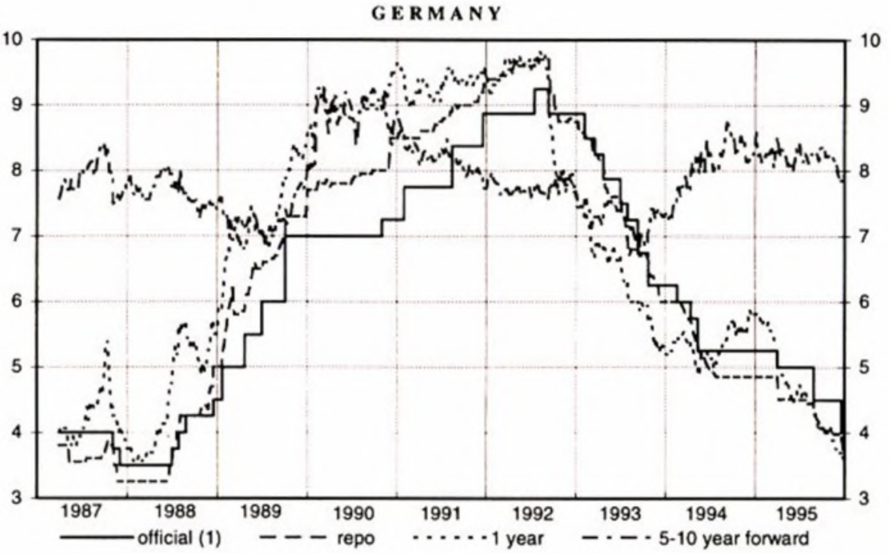
21 The Bank of Italy started to tighten monetary conditions through repo operations in June 1994 before the turning point in consumer prices, reacting to rising producer prices and deteriorating expectations emerging both from surveys and other real and financial indicators (see Banca d'Italia, 1995).

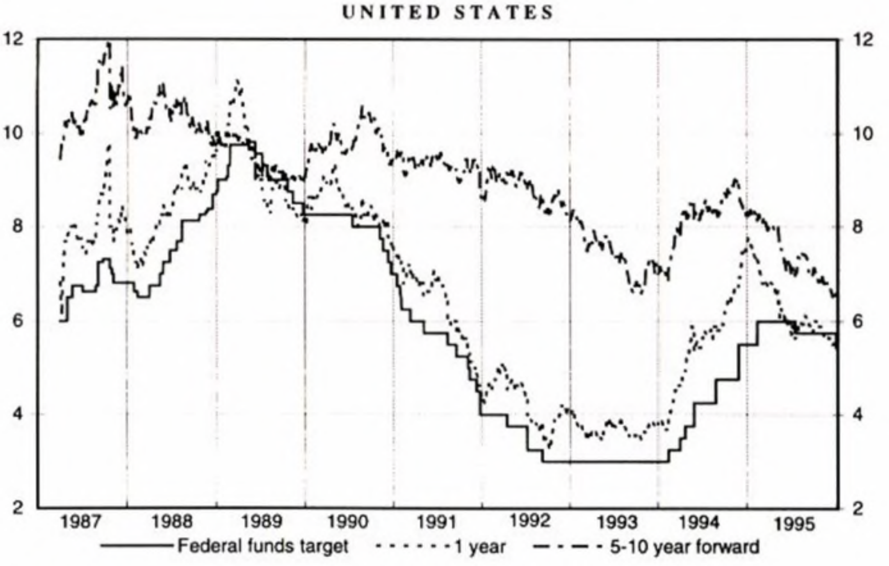
1995 - when official rates were again raised - as the coefficient shows a sizable reduction possibly reflecting an improved market perception concerning the determination of the Bank of Italy in fighting inflation and the long-run sustainability of the policy stance.

In Germany the coefficient shows substantial stability (Figure 3c), even when estimated with $\chi=0.1$, remaining negative during the whole period. Interestingly, during the phase of tightening following reunification, the impact effect of official rate rises on long-term forward rates was a reduction of long-term forward rates. This suggests that these moves were perceived by the market as effectively signalling the Bundesbank's resolve to maintain an anti-inflationary policy even in the face of dramatic shocks. On the other hand, the impact effect of more recent cuts, though undertaken during an economic downturn, seems to have been a slight increase of long-term inflation expectations, as proxied by long-term forward rates.

Fig. 1a

MONETARY POLICY RATES AND MARKET RATES
(end-of-week data; percentage values)



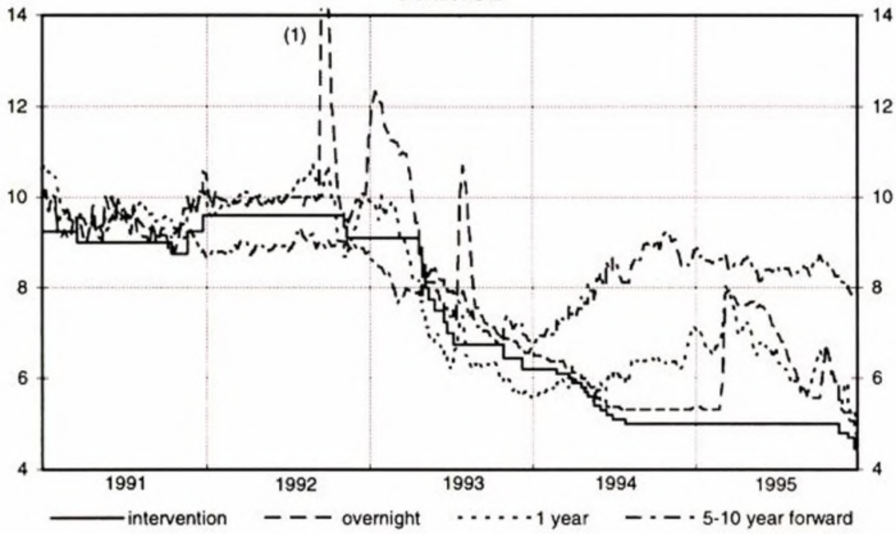


(1) Average of the discount rate and of the Lombard rate.

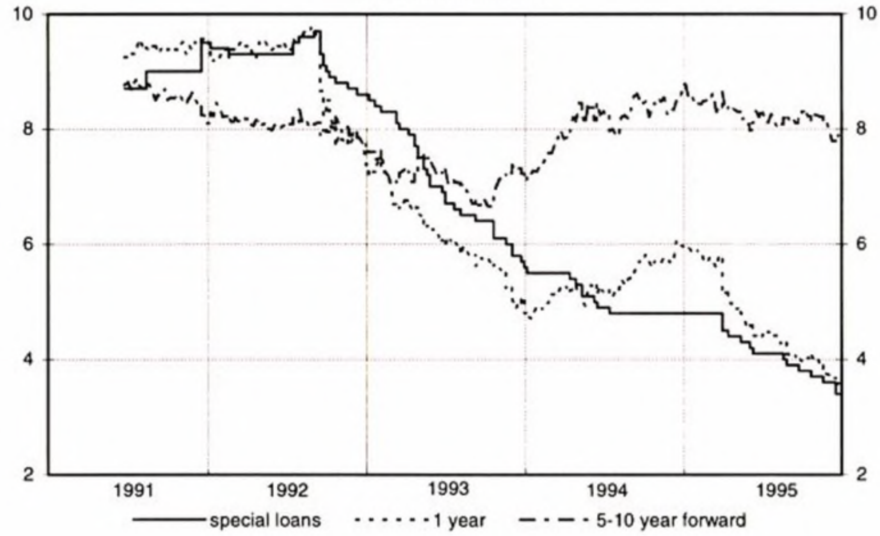
Fig. 1b

MONETARY POLICY RATES AND MARKET RATES
(end-of-week data; percentage values)

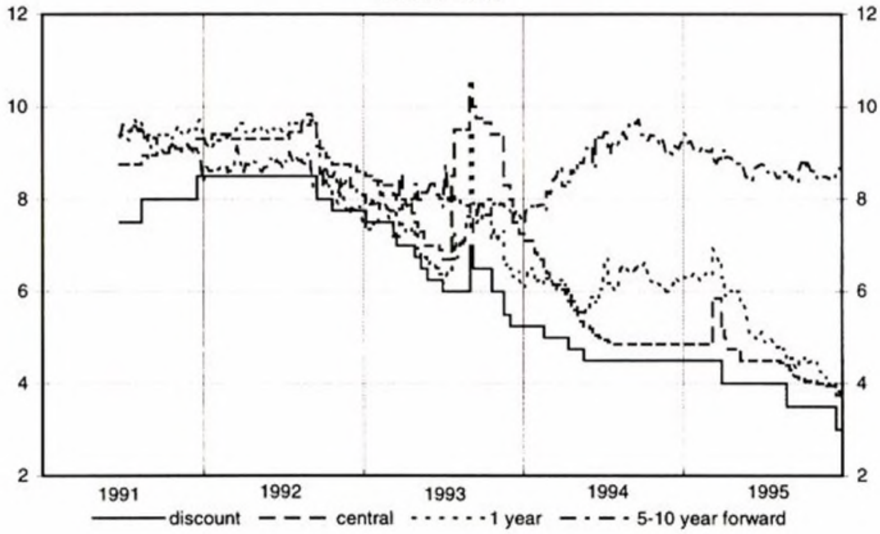
FRANCE



NETHERLANDS



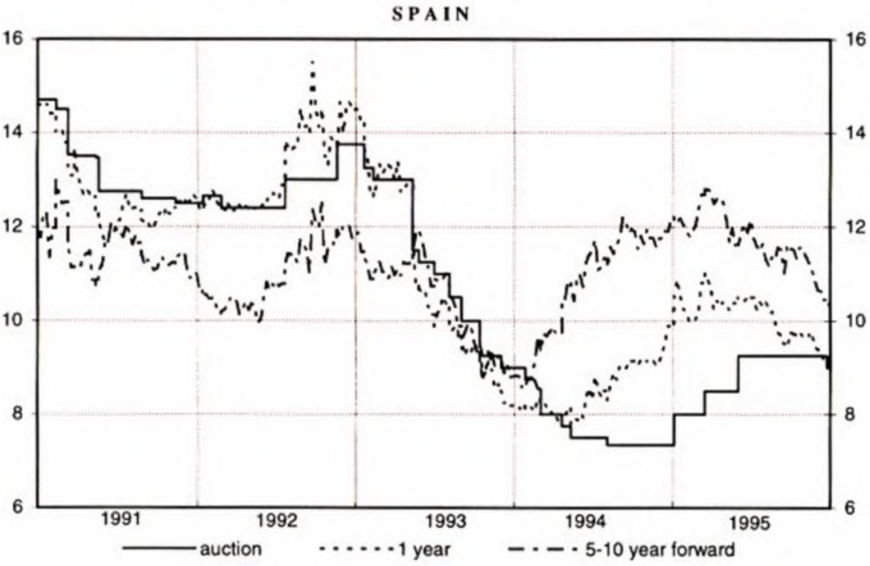
BELGIUM



(1) Maximum = 19.88.

Fig. 1c

MONETARY POLICY RATES AND MARKET RATES
(end-of-week data; percentage values)



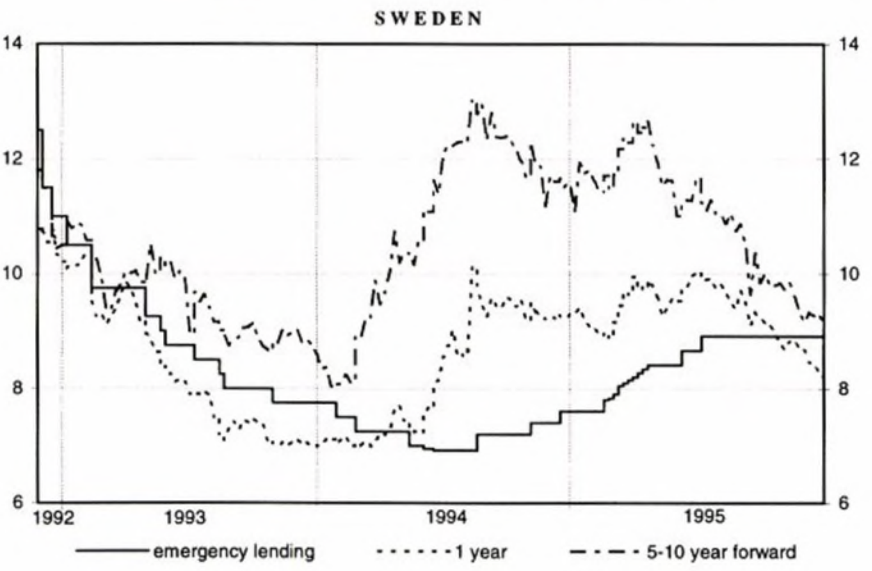
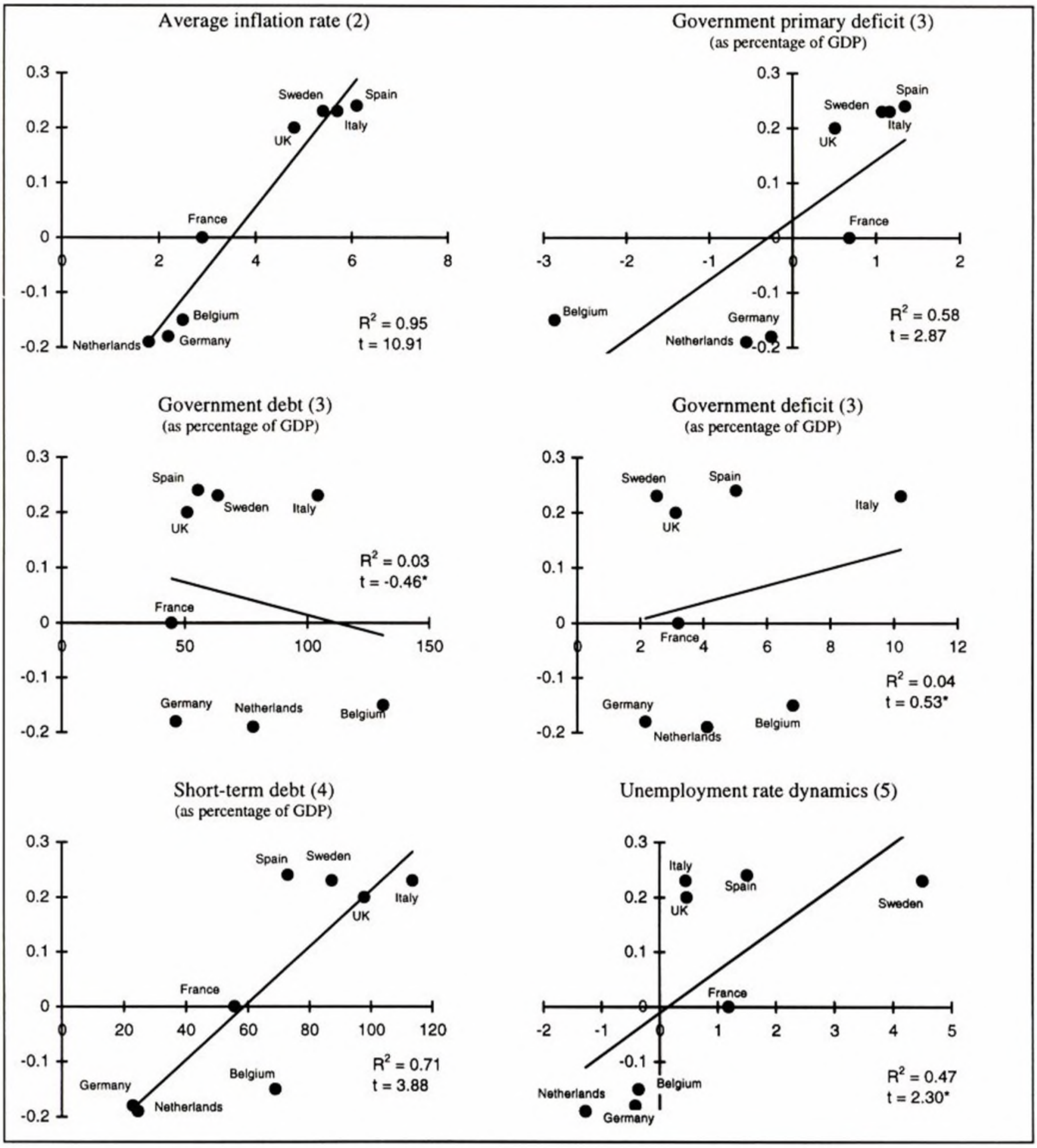


Fig. 2a

EFFECT OF MONETARY POLICY RATE CHANGES ON LONG-TERM FORWARD RATES VIS-À-VIS MACROECONOMIC VARIABLES (1)



Sources: Bank of Italy, BIS, CSO, OECD.

(1) On the vertical axis: coefficients of policy rates changes in the 5-10 year forward rate regression (see table 1); on the horizontal axis: macroeconomic variables. A star by the *t*-value indicates that the slope coefficient of the interpolating line is not significant at the 5% confidence level.

(2) Annual (geometric) average rates over the 1985-1995 period.

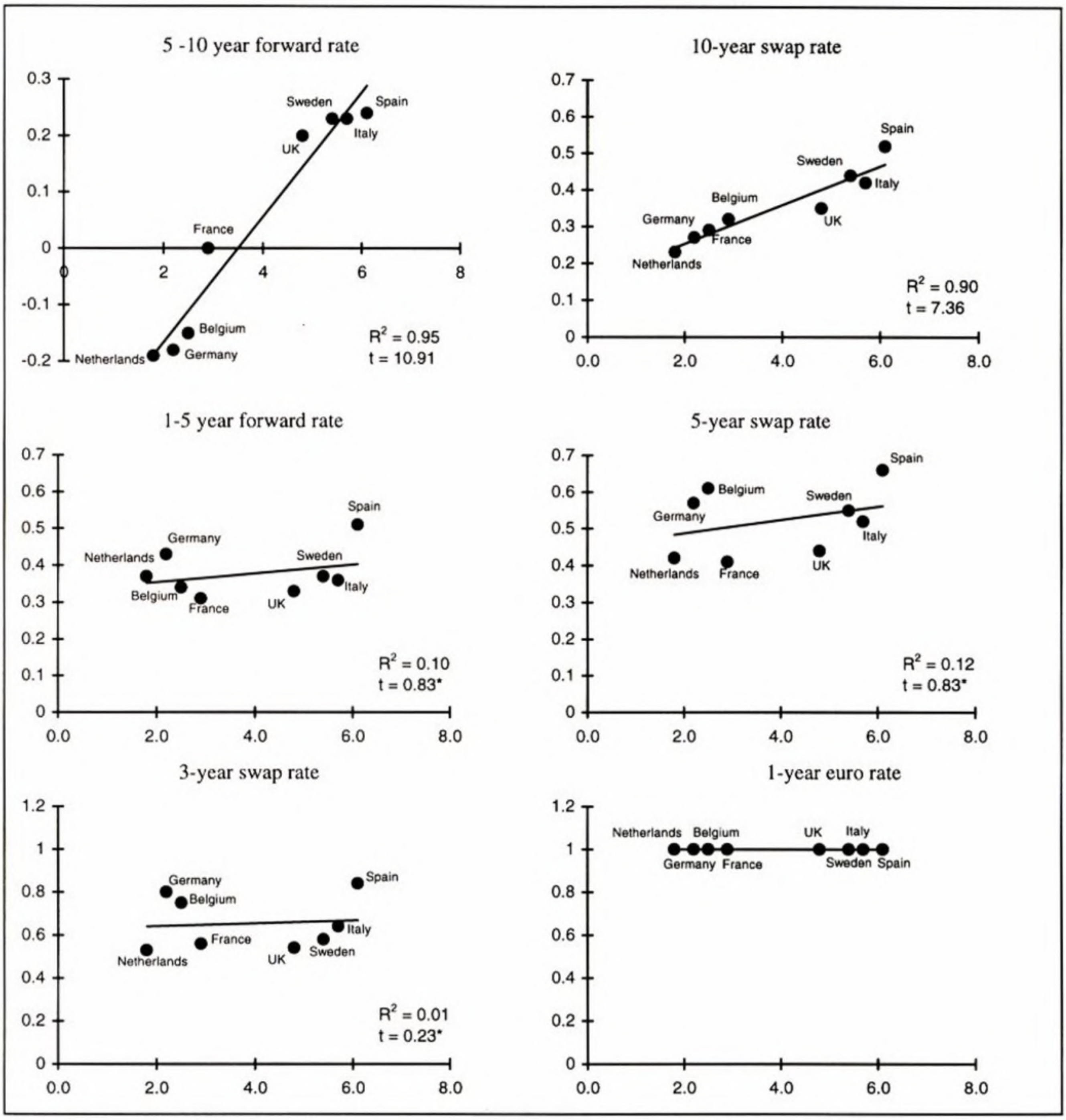
(3) Annual averages over the periods indicated in footnote 2.

(4) Total of short-term (up to 1 year) and variable interest rate credit to both the private and the public sectors as percentage of GDP, measured in 1992 for all countries but Belgium (1990) and Sweden (1991).

(5) Difference between the averages over the 1991-1995 and the 1985-1990 periods.

Fig. 2b

EFFECT OF MONETARY POLICY RATE CHANGES ON THE TERM STRUCTURE VIS-À-VIS THE AVERAGE INFLATION RATE (1)

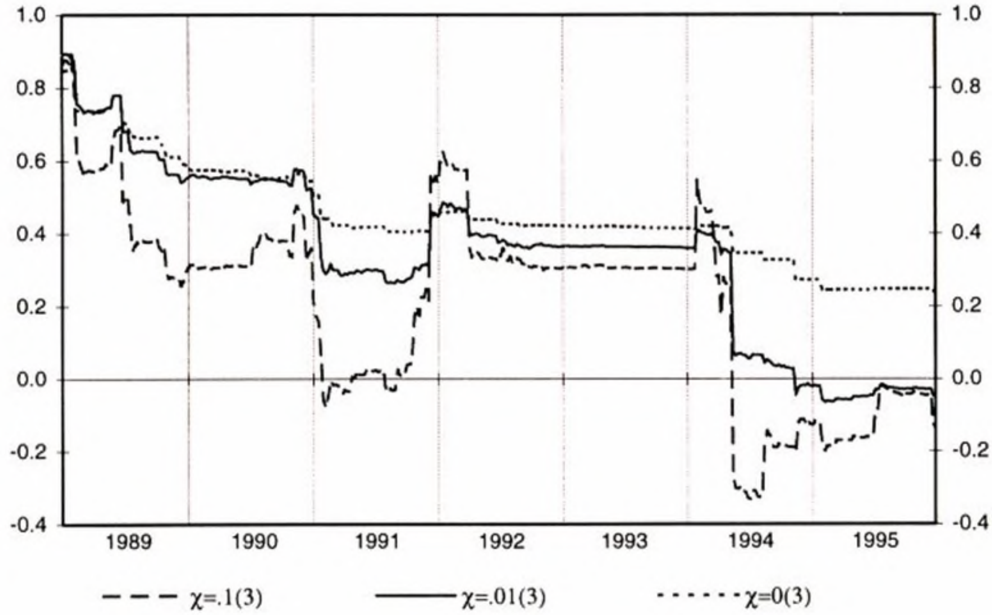


Source: OECD, Economic Outlook.

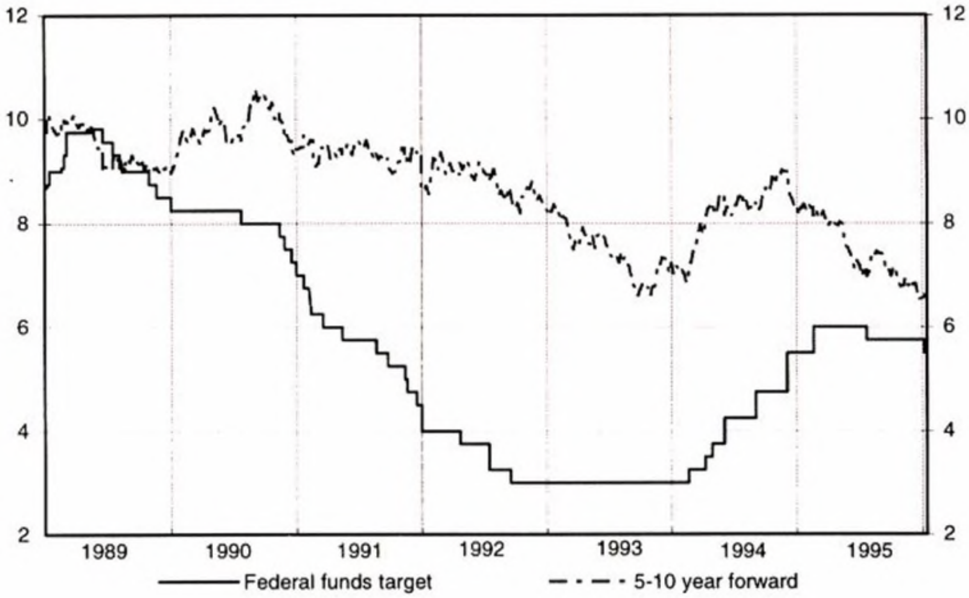
(1) On the vertical axis: coefficients of policy rate changes in the indicated market rates regressions (see table 1); on the horizontal axis: annual (geometric) average inflation rate over the 1985-1995 period. A star by the t-value indicates that the slope coefficient of the interpolating line is not significant at the 5% confidence level.

UNITED STATES
(end-of-week data)

Kalman filter: monetary policy rate[(1)] coefficient in the 5-10 year forward rate regression[(2)]



Monetary policy rate and 5-10 year forward rate

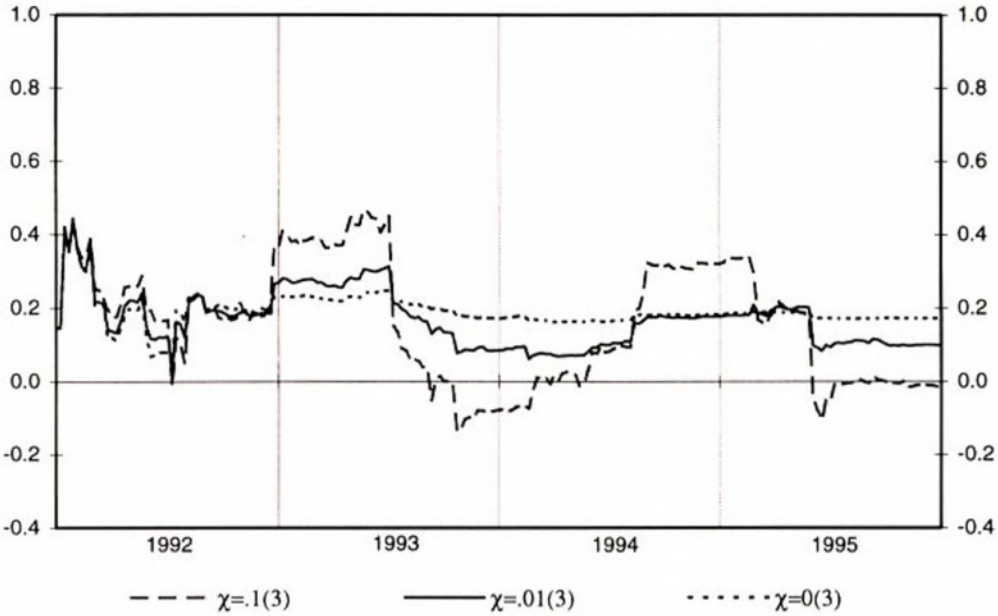


(1) Federal funds target.
(2) Coefficient of the contemporaneous value of the change in the Federal funds rate.
(3) The parameter χ represents the signal-to-noise ratio (see equation 12 in the text). The case of χ=0 corresponds to a recursive OLS estimation.

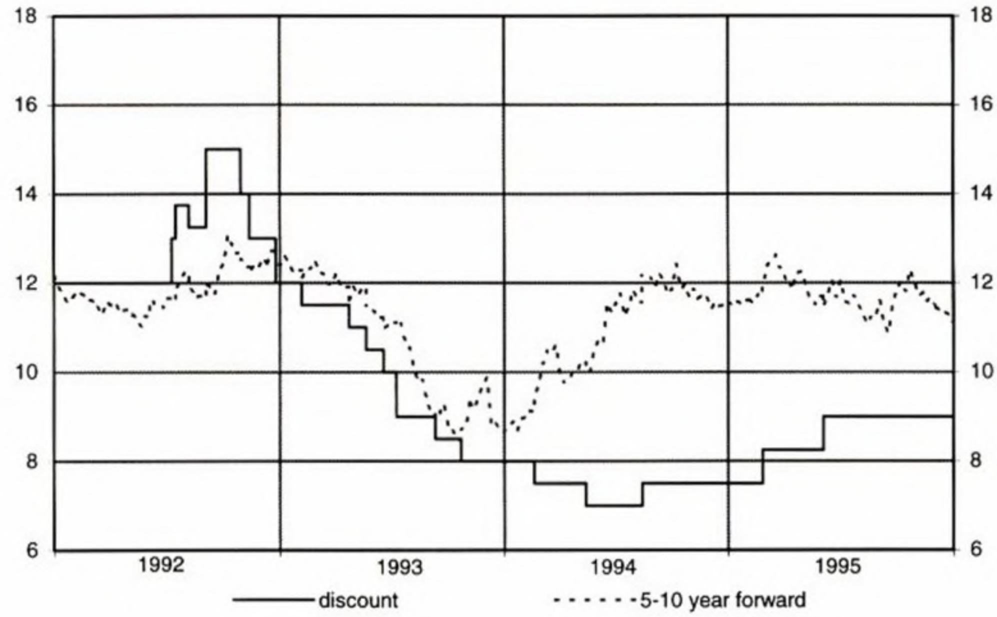
Fig. 3b

ITALY
(end-of-week data)

Kalman filter: monetary policy rate(1) coefficient in the 5-10 year forward rate regression(2)



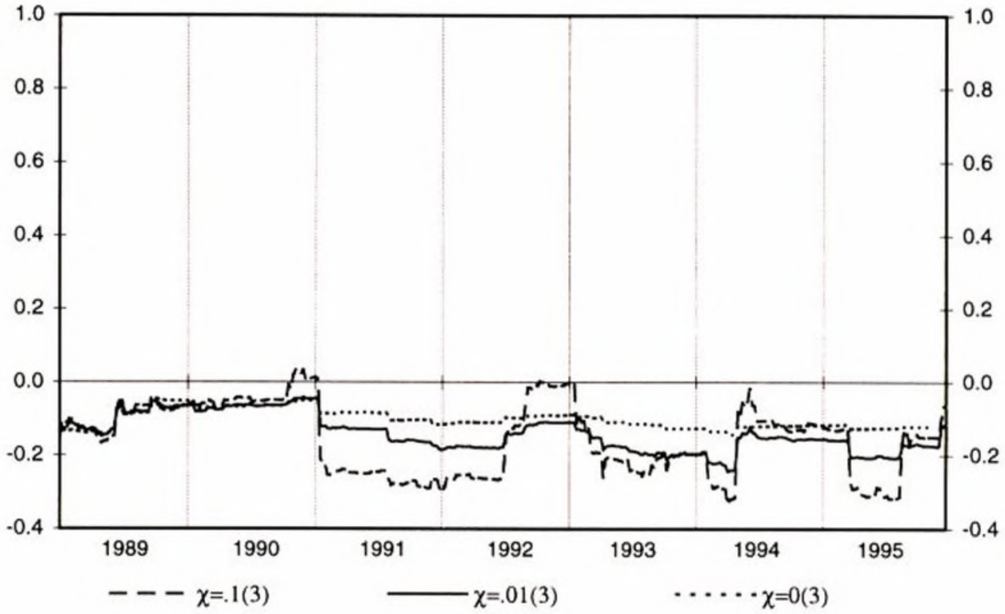
Monetary policy rate and 5-10 year forward rate



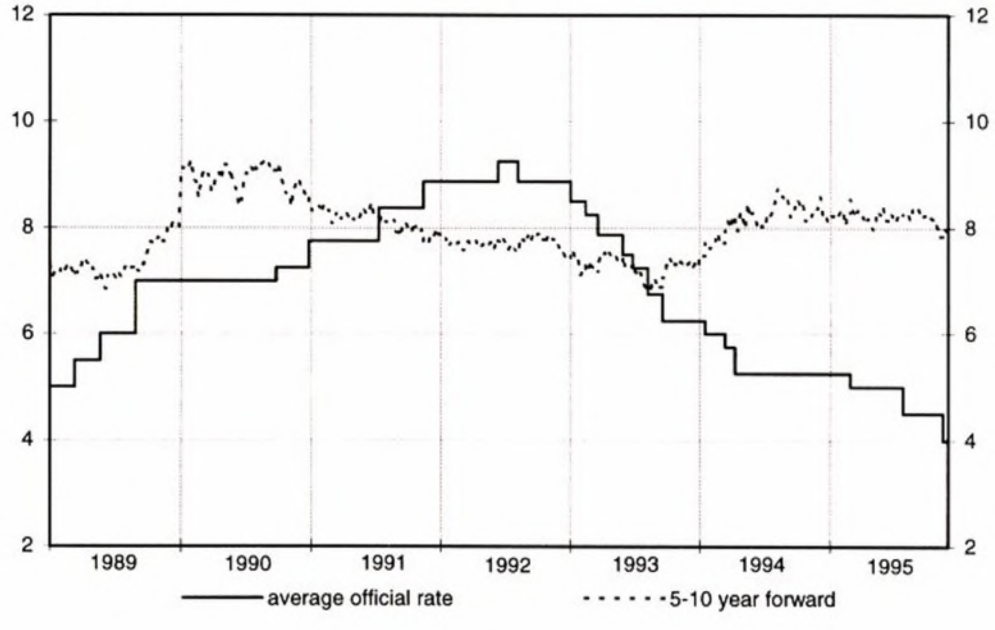
(1) Discount rate.
(2) Coefficient of the contemporaneous value of the change in the average of the discount rate.
(3) The parameter χ represents the signal-to-noise ratio (see equation 12 in the text). The case of χ=0 corresponds to a recursive OLS estimation.

GERMANY
(end-of-week data)

Kalman filter: monetary policy rate[1] coefficient in the 5-10 year forward rate regression[2]



Monetary policy rate and 5-10 year forward rate



(1) Average of the discount rate and the Lombard rate.

(2) Coefficient of the contemporaneous value of the change in the average of the discount rate and the Lombard rate.

(3) The parameter χ represents the signal-to-noise ratio (see equation 12 in the text). The case of χ=0 corresponds to a recursive OLS estimation.

Table 1

EFFE CTS OF POLICY RATE CHANGES ON THE TERM STRUCTURE (1)

Endogenous variable	Germany	Netherlands	Belgium	Spain	Sweden	Italy	UK	France	USA 1987-95	USA 1991-95
Ytm (2)										
1 year (3)	1.00 -	1.00 -	1.00 -	1.00 -	1.00 -	1.00 -	1.00 -	1.00 -	1.00 -	1.00 -
3 years	0.80 (4.68)	0.53 (2.65)	0.75 (3.38)	0.84 (5.10)	0.58 (5.81)	0.64 (6.21)	0.54 (6.32)	0.56 (4.84)	0.70 (2.65)	0.46 (1.76)
5 years	0.57 (4.40)	0.42 (2.08)	0.61 (3.28)	0.66 (4.92)	0.55 (8.54)	0.52 (5.52)	0.44 (6.67)	0.41 (3.35)	0.60 (2.56)	0.29 (1.29)
10 years	0.27 (3.14)	0.23 (2.03)	0.29 (2.61)	0.52 (5.30)	0.44 (4.67)	0.42 (4.95)	0.35 (6.14)	0.32 (2.97)	0.49 (1.99)	0.16 (0.70)
Forward (4)										
1-5 years	0.43 (3.79)	0.37 (2.36)	0.34 (2.25)	0.51 (3.96)	0.37 (6.52)	0.36 (3.48)	0.33 (5.30)	0.31 (2.34)	0.50 (2.19)	0.16 (0.73)
5-10 years	-0.18 (-2.17)	-0.19 (-1.96)	-0.14 (-4.68)	0.24 (2.53)	0.23 (2.11)	0.23 (3.31)	0.20 (15.90)	0.00 (-0.04)	0.35 (1.28)	-0.01 (-0.03)

(1) OLS estimates; sample periods, number of policy rates and number of lags included in the regressions are indicated in Tables 2a-2i. Estimates refer to the effect of policy rate increases for France and the United Kingdom and of all changes for the other countries. F-statistics, based on Newey and West (1987) standard errors in parentheses. - (2) Yields to maturity: 1-year Libor and 3-year, 5-year and 10-year swap rates. - (3) The restriction of a unitary coefficient for the policy rates changes in the 1-year rate regression is accepted in all countries. - (4) 1-year ahead 4-year forward rate and 5-year ahead 5-year forward rate.

Table 2a

GERMANY:
EFFECTS OF CHANGES IN MONETARY POLICY RATES ON THE TERM STRUCTURE (1)

Endogenous variable	Δi^{P1}	Δi^{P2}	$\Delta i^{P1} + \Delta i^{P2}$	Δi^{*}	Δi	α_0	*rsc(2)*	*Wald(3)*
Ytm (4)								
1 year	-0.07	1.17	1.10	0.25	-0.14	0.00	0.38	-
	(-0.87)	(7.45)	(7.44)	(3.95)	(-2.83)	(-0.29)		
	-0.10	*1.10*	*1.00*	*0.24*	*-0.12*	*0.00*	-	*0.50*
	(-1.08)	*(11.62)*		*(5.71)*	*(-2.87)*	*(-0.10)*		
3 years	-0.03	0.83	0.80	0.28	0.04	0.00	0.30	-
	(-0.42)	(5.06)	(4.68)	(4.62)	(0.73)	(0.23)		
5 years	-0.04	0.61	0.57	0.31	0.05	0.00	0.72	-
	(-0.72)	(5.50)	(4.40)	(7.89)	(1.11)	(0.50)		
10 years	-0.07	0.35	0.27	0.30	0.04	0.00	0.88	-
	(-3.31)	(4.93)	(3.14)	(6.84)	(0.73)	(0.55)		
Forward (5)								
1-5 years	-0.03	0.46	0.43	0.26	-0.02	0.00	0.65	-
	(-0.64)	(4.31)	(3.79)	(6.74)	(-0.48)	(0.52)		
5-10 years	-0.14	-0.04	-0.18	0.23	-0.10	0.00	0.32	-
	(-3.46)	(-0.57)	(-2.17)	(3.99)	(-1.33)	(0.54)		

(1) OLS estimates; sample period: April 1987-December 1995. The estimates refer to the sum of the coefficients for each variable; in all regressions one lead is included for the policy rates and one lag for all variables. F statistics, based on Newey and West (1987) standard errors in parentheses. - (2) P-prob of the Lagrange multiplier test of residual serial correlation. - (3) P-prob of the Wald test of the restriction that the total effect of the policy rates changes equals 1. - (4) Yields to maturity: 1-year Libor and 3-year, 5-year and 10-year swap rates. - (5) 1-year ahead 4-year forward rate and 5-year ahead 5-year forward rate.

Table 2b

THE NETHERLANDS:
EFFECTS OF CHANGES IN MONETARY POLICY RATES ON THE TERM STRUCTURE (1)

Endogenous variable	Δi^{P1}	Δi^{P2}	$\Delta i^{P1} + \Delta i^{P2}$	Δi^{*}	Δi	α_0	*rsc(2)*	*Wald(3)*
Ytm (4)								
1 year	1.05 (3.56)	-	-	0.29 (3.91)	-0.18 (-5.46)	0.00 (-0.53)	0.44	-
	1.00	-	-	*0.28* *(3.30)*	*-0.16* *(-2.00)*	*0.00* *(-0.53)*	-	*0.85*
3 years	0.53 (2.65)	-	-	0.20 (3.65)	0.18 (3.08)	0.00 (-0.31)	0.10	-
5 years	0.42 (2.08)	-	-	0.26 (5.62)	0.17 (2.63)	0.00 (-0.10)	0.62	-
10 years	0.23 (2.03)	-	-	0.31 (9.14)	0.04 (1.08)	0.00 (-0.28)	0.25	-
Forward (5)								
1-5 years	0.37 (2.36)	-	-	0.28 (6.26)	0.09 (1.32)	0.00 (0.03)	0.87	-
5-10 years	-0.19 (-1.96)	-	-	0.41 (11.18)	-0.37 (-5.39)	0.00 (-0.65)	0.92	-

(1) OLS estimates; sample period: June 1991-December 1995. The estimates refer to the sum of the coefficients for each variable; in all regressions one lead is included for the policy rate and two lags for all variables. F statistics, based on Newey and West (1987) standard errors in parentheses. - (2) P-prob of the Lagrange multiplier test of residual serial correlation. - (3) P-prob of the Wald test of the restriction that the total effect of the policy rate changes equals 1. - (4) Yields to maturity: 1-year Libor and 3-year, 5-year and 10-year swap rates. - (5) 1-year ahead 4-year forward rate and 5-year ahead 5-year forward rate.

Table 2c

BELGIUM:
EFFECTS OF CHANGES IN MONETARY POLICY RATES ON THE TERM STRUCTURE (1)

Endogenous variable	Δi^{P1}	Δi^{P2}	$\Delta i^{P1} + \Delta i^{P2}$	Δi^{*}	Δi	α_0	*rsc(2)*	*Wald(3)*
Ytm (4)								
1 year	0.91	0.35	1.27	0.20	-0.23	0.00	0.11	-
	(2.19)	(3.15)	(3.63)	(2.21)	(-3.10)	(-0.19)		
	0.67	*0.33*	*1.00*	*0.17*	*-0.18*	*-0.01*	-	*0.59*
	(5.59)	*(2.77)*		*(1.13)*	*(-2.07)*	*(-0.49)*		
3 years	0.61	0.14	0.75	0.18	-	0.00	0.60	-
	(2.32)	(2.77)	(3.38)	(3.86)		(-0.22)		
5 years	0.52	0.09	0.61	0.23	-	0.00	0.82	-
	(2.37)	(1.88)	(3.28)	(4.40)		(-0.06)		
10 years	0.29	0.00	0.29	0.28	-	0.00	0.45	-
	(2.05)	(0.00)	(2.61)	(7.01)		(-0.23)		
Forward (5)								
1-5 years	0.33	0.01	0.34	0.22	-	0.00	0.96	-
	(1.87)	(0.33)	(2.25)	(3.85)		(-0.27)		
5-10 years	-	-0.14	0.14	0.23	-0.58	0.00	0.15	-
		(-4.68)	(-4.68)	(4.25)	(-6.67)	(-0.33)		

(1) OLS estimates; sample period: June 1991-December 1995. The estimates refer to the sum of the coefficients for each variable; in all regressions one lead is included for the policy rates; two lags are included for all variables in the regression for the 1-year rate; no lags are included in the other regressions, with the exception of that for the 5-10 year forward rate, in which two lags are included for the endogenous variable. F statistics, based on Newey and West (1987) standard errors in parentheses. - (2) P-prob of the Lagrange multiplier test of residual serial correlation. - (3) P-prob of the Wald test of the restriction that the total effect of the policy rates changes equals 1. - (4) Yields to maturity: 1-year Libor and 3-year, 5-year and 10-year swap rates. - (5) 1-year ahead 4-year forward rate and 5-year ahead 5-year forward rate.

Table 2d

SPAIN:
EFFECTS OF CHANGES IN MONETARY POLICY RATES ON THE TERM STRUCTURE (1)

Endogenous variable	Δi^{P1}	Δi^{P2}	$\Delta i^{P1} + \Delta i^{P2}$	Δi^{*}	Δi	α_0	*rsc(2)*	*Wald(3)*
Ytm (4)								
1 year	1.29 (6.38)	-	-	0.29 (2.77)	-0.23 (-2.56)	0.00 (0.24)	0.73	-
	1.00	-	-	*0.28* *(1.84)*	*-0.20* *(-3.52)*	*0.00* *(-0.14)*	-	*0.07*
3 years	0.84 (5.10)	-	-	0.13 (0.49)	-0.11 (-1.04)	0.00 (-0.06)	0.62	-
5 years	0.66 (4.92)	-	-	0.34 (1.54)	-0.05 (-0.56)	0.00 (0.00)	0.88	-
10 years	0.52 (5.30)	-	-	0.53 (3.06)	-0.04 (-0.65)	0.00 (0.29)	0.33	-
Forward (5)								
1-5 years	0.51 (3.96)	-	-	0.33 (1.43)	-0.10 (-1.16)	0.00 (0.10)	0.73	-
5-10 years	0.24 (2.53)	-	-	0.40 (1.77)	-0.07 (-1.72)	0.00 (0.30)	0.94	-

(1) OLS estimates; sample period: January 1991-December 1995. The estimates refer to the sum of the coefficients for each variable; in all regressions one lead is included for the policy rate and one lag for all variables. F statistics, based on Newey and West (1987) standard errors in parentheses. - (2) P-prob of the Lagrange multiplier test of residual serial correlation. - (3) P-prob of the Wald test of the restriction that the total effect of the policy rate changes equals 1. - (4) Yields to maturity: 1-year Libor and 3-year, 5-year and 10-year swap rates. - (5) 1-year ahead 4-year forward rate and 5-year ahead 5-year forward rate.

Table 2e

SWEDEN:
EFFECTS OF CHANGES IN MONETARY POLICY RATES ON THE TERM STRUCTURE (1)

Endogenous variable	Δi^{P1}	Δi^{P2}	$\Delta i^{P1} + \Delta i^{P2}$	Δi^{*}	Δi	α_0	*rsc(2)*	*Wald(3)*
Ytm (4)								
1 year	1.29 (10.51)	-	-	0.38 (3.69)	-0.20 (-3.22)	0.00 (-0.22)	0.36	-
	1.00	-	-	*0.35* *(1.67)*	*-0.12* *(-1.37)*	*-0.01* *(-0.50)*	-	*0.38*
3 years	0.58 (5.81)	-	-	0.42 (5.04)	-0.03 (-0.35)	0.00 (-0.12)	0.62	-
5 years	0.55 (8.54)	-	-	0.60 (4.54)	-0.02 (-0.19)	0.00 (-0.08)	0.92	-
10 years	0.44 (4.67)	-	-	0.78 (3.26)	-0.07 (-0.58)	0.00 (0.00)	0.82	-
Forward (5)								
1-5 years	0.37 (6.52)	-	-	0.64 (4.93)	-0.02 (-0.20)	0.00 (0.04)	0.92	-
5-10 years	0.23 (2.11)	-	-	1.12 (4.29)	-0.31 (-2.21)	0.00 (0.19)	0.90	-

(1) OLS estimates; sample period: November 1992-December 1995. The estimates refer to the sum of the coefficients for each variable; in all regressions one lead is included for the policy rate and two lags for all variables. F statistics, based on Newey and West (1987) standard errors in parentheses. - (2) P-prob of the Lagrange multiplier test of residual serial correlation. - (3) P-prob of the Wald test of the restriction that the total effect of the policy rate changes equals 1. - (4) Yields to maturity: 1-year Libor and 3-year, 5-year and 10-year swap rates. - (5) 1-year ahead 4-year forward rate and 5-year ahead 5-year forward rate.

Table 2f

ITALY:
EFFECTS OF CHANGES IN MONETARY POLICY RATES ON THE TERM STRUCTURE (1)

Endogenous variable	Δi^{P1}	Δi^{P2}	$\Delta i^{P1} + \Delta i^{P2}$	Δi^{*}	Δi	α_0	$rsc(2)$	$Wald(3)$
Ytm (4)								
1 year	0.57 (4.11)	0.62 (12.16)	1.19 (9.34)	0.25 (1.19)	-0.42 (-2.85)	0.00 (0.06)	0.07	-
	0.38 *(4.78)*	*0.62* *(7.91)*	*1.00*	*0.19* *(0.95)*	*-0.40* *(-6.94)*	*-0.00* *(-0.10)*	-	*0.09*
3 years	0.50 (4.71)	0.15 (2.47)	0.64 (6.21)	0.29 (1.86)	-0.06 (-1.28)	0.00 (0.05)	0.19	-
5 years	0.42 (4.79)	0.11 (2.06)	0.52 (5.52)	0.50 (4.98)	-0.02 (-0.37)	0.00 (0.21)	0.44	-
10 years	0.36 (4.29)	0.05 (0.95)	0.42 (4.95)	0.53 (4.64)	-0.03 (-0.53)	0.00 (0.24)	0.97	-
Forward (5)								
1-5 years	0.40 (5.27)	-0.03 (-0.55)	0.36 (3.48)	0.50 (6.07)	-	0.00 (0.18)	0.73	-
5-10 years	0.27 (3.02)	-0.04 (-0.56)	0.23 (3.31)	0.41 (2.22)	-0.06 (-1.06)	0.00 (0.08)	0.95	-

(1) OLS estimates; sample period: March 1991-December 1995. The estimates refer to the sum of the coefficients for each variable; in all regressions one lead is included for the policy rates; one lag is included for all variables in all regressions, with the exception of that for the 1-5 year forward rate, in which no lags are included for the endogenous variable. F statistics, based on Newey and West (1987) standard errors in parentheses. - (2) P-prob of the Lagrange multiplier test of residual serial correlation. - (3) P-prob of the Wald test of the restriction that the total effect of the policy rates changes equals 1. - (4) Yields to maturity: 1-year Libor and 3-year, 5-year and 10-year swap rates. - (5) 1-year ahead 4-year forward rate and 5-year ahead 5-year forward rate.

Table 2g

UNITED KINGDOM:
EFFECTS OF CHANGES IN MONETARY POLICY RATES ON THE TERM STRUCTURE (1)

Endogenous variable	$\Delta i^{P\uparrow}$	$\Delta i^{P\downarrow}$	Δi^{*}	Δi	α_0	*rsc(2)*	*Wald(3)*
Ytm (4)							
1 year	0.93 (8.28)	1.13 (10.08)	0.30 (7.46)	-0.18 (-4.96)	0.00 (0.66)	0.41	0.16
	1.00	*1.13* *(8.67)*	*0.30* *(4.02)*	*-0.18* *(-4.14)*	*0.00* *(0.30)*	-	*0.48*
3 years	0.54 (6.32)	0.73 (4.92)	0.35 (4.65)	-0.06 (-1.02)	0.00 (0.60)	0.27	0.27
5 years	0.44 (6.67)	0.55 (5.45)	0.34 (6.13)	-0.01 (-0.22)	0.00 (0.43)	0.30	0.35
10 years	0.35 (6.14)	0.31 (7.80)	0.43 (5.94)	-0.03 (-0.55)	0.00 (-0.09)	0.89	0.38
Forward (5)							
1-5 years	0.33 (5.30)	0.44 (4.17)	0.35 (5.07)	-0.08 (-1.77)	0.00 (0.37)	0.18	0.33
5-10 years	0.20 (15.90)	-0.10 (-0.74)	0.47 (4.09)	-0.13 (-1.64)	0.00 (-0.80)	0.60	0.04

(1) OLS estimates; sample period: April 1987-December 1995. The estimates refer to the sum of the coefficients for each variable; in all regressions one lead is included for the policy rate and one lag for all variables. F statistics, based on Newey and West (1987) standard errors in parentheses. - (2) P-prob of the Lagrange multiplier test of residual serial correlation. - (3) For the 1-year rate P-prob of the Wald test of the restriction that the total effect of the policy rate changes equals 1; for other rates of the restriction that the effects of policy rate increases and reductions are equal. - (4) Yields to maturity: 1-year Libor and 3-year, 5-year and 10-year swap rates. - (5) 1-year ahead 4-year forward rate and 5-year ahead 5-year forward rate.

Table 2h

FRANCE:
EFFECTS OF CHANGES IN MONETARY POLICY RATES ON THE TERM STRUCTURE (1)

Endogenous variable	$\Delta i^{P1\uparrow} + \Delta i^{P2\uparrow}$	$\Delta i^{P1\downarrow} + \Delta i^{P2\downarrow}$	Δi^{*}	Δi	α_0	*rsc(2)*	*Wald(3)*
<u>*Ytm*</u> (4)							
1 year	0.89 (4.04)	0.76 (5.66)	0.19 (1.90)	0.00 (0.03)	0.00 (0.02)	0.93	-
	1.00	*0.76* *(10.23)*	*0.20* *(1.64)*	*0.00* *(0.02)*	*0.00* *(-0.01)*	-	0.62
3 years	0.56 (4.84)	0.42 (6.04)	0.17 (1.83)	0.00 (-0.14)	0.00 (0.13)	0.60	0.29
5 years	0.41 (3.35)	0.29 (4.08)	0.24 (3.06)	0.07 (1.85)	0.00 (0.16)	0.18	0.43
10 years	0.32 (2.97)	0.15 (1.57)	0.35 (4.46)	-0.04 (-1.01)	0.00 (0.24)	0.59	0.31
<u>*Forward*</u> (5)							
1-5 years	0.31 (2.34)	0.17 (1.64)	0.27 (3.03)	0.02 (0.48)	0.00 (0.20)	0.99	0.47
5-10 years	0.00 (-0.04)	-0.14 (-1.02)	0.41 (3.28)	-0.22 (-6.12)	0.00 (0.53)	0.16	0.48

(1) OLS estimates; sample period: January 1991-December 1995. The estimates refer to the sum of the coefficients for each variable; in all regressions one lead is included for the policy rates and one lag for all variables. F statistics, based on Newey and West (1987) standard errors in parentheses. - (2) P-prob of the Lagrange multiplier test of residual serial correlation. - (3) For the 1-year rate P-prob of the Wald test of the restriction that the total effect of the policy rates increases equals 1; for other rates of the restriction that the effects of policy rates increases and reductions are equal. - (4) Yields to maturity: 1-year Libor and 3-year, 5-year and 10-year swap rates. - (5) 1-year ahead 4-year forward rate and 5-year ahead 5-year forward rate.

Table 2i

UNITED STATES:
EFFECTS OF CHANGES IN MONETARY POLICY RATES ON THE TERM STRUCTURE (1)

Endogenous variable	Δi^{P1}	Δi^{P2}	$\Delta i^{P1} + \Delta i^{P2}$	Δi^{*}	Δi	α_0	*rsc(2)*	*Wald(3)*
Ytm (4)								
1 year	1.04 (3.06)	-	-	-	-0.15 (-0.99)	0.00 (-0.76)	0.39	-
	1.00	-	-	-	*-0.16* *(-3.59)*	*0.00* *(-0.11)*	-	*0.75*
3 years	0.70 (2.65)	-	-	-	-0.13 (-2.28)	0.00 (-0.73)	0.93	-
5 years	0.60 (2.56)	-	-	-	-0.14 (-2.24)	0.00 (-1.02)	0.95	-
10 years	0.49 (1.99)	-	-	-	-0.14 (-2.01)	0.00 (-0.99)	0.69	-
Forward (5)								
1-5 years	0.50 (2.19)	-	-	-	-0.21 (-3.58)	0.00 (-0.88)	0.53	-
5-10 years	0.35 (1.28)	-	-	-	-0.15 (-1.94)	0.00 (-0.90)	0.34	-

(1) OLS estimates; sample period: April 1987-December 1995. The estimates refer to the sum of the coefficients for each variable; in all regressions one lead is included for the policy rate and one lag for all variables. F statistics, based on Newey and West (1987) standard errors in parentheses. - (2) P-prob of the Lagrange multiplier test of residual serial correlation. - (3) P-prob of the Wald test of the restriction that the total effect of the policy rate changes equals 1. - (4) Yields to maturity: 1-year Libor and 3-year, 5-year and 10-year swap rates. - (5) 1-year ahead 4-year forward rate and 5-year ahead 5-year forward rate.

APPENDIX

Derivation of the forward rates from the swap curve

Forward rates were derived from swap rates according to the following simple procedure (see Drudi and Giraldi, 1991).

Let us define i(j) as the (not compounded) interest rate on a zero coupon with maturity j years from now. Also define the discount factor P_j as:

$$P_j = 1/(1+i(j)j)$$

Given that the swap rate is that rate which equates a floating rate flow with a fixed rate flow, the following equivalence holds in the case of a 2-year swap rate:

$$sr_2 P_1 + sr_2 P_2 = i(1) P_1 + E(i(1,2)) P_2$$

where sr_2 is the 2-year swap rate and E(i(1,2)) is the expected value (adjusted for risk) of a one-year ahead one-year rate. Adding $100P_2$ on both sides we have:

$$sr_2 P_1 + sr_2 P_2 + 100 P_2 = i(1) P_1 + E(i(1,2)) P_2 + 100 P_2 .$$

The right-hand side is the cash flow of a floater that, in a default-free world, has a price of approximately 100. Thus we can write:

$$100 = sr_2 P_1 + sr_2 P_2 + 100 P_2 .$$

Given P_1 (from the one-year euro-deposit rate, which is a zero coupon rate) P_2 can be deduced. The forward rate can then be obtained from P_1 and P_2 as:

$$f(1,2) = (P_1 / P_2 - 1)100 .$$

Given P_1 , P_2 and the 3-year swap rate the same procedure can be applied to derive P_3 and thus to obtain the 2-year ahead 1-year forward rate and so on.

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