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Measuring the long-term perception of monetary policy and the term structure

Suomen Pankin keskustelualoitteita
Finlands Banks diskussionsunderlag

Suomen Pankki
Bank of Finland
P.O.Box 160
FIN-00101 HELSINKI
Finland
☎ + 358 9 1831

http://www.bof.fi

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Measuring the long-term perception of monetary policy and the term structure

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Nicolas Rautureau
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Abstract

This paper has two objectives. The first is to identify the long-term public perception of monetary policy. The second is to identify the relationship between this perception and long-term bond rates. For German data, the use of a two-factor model of the term structure results in the best forecast of long-term interest rates for the period between January 1975 and January 2003. It also allows us to introduce as the second factor the long-term perception of inflation as a characteristic of the behaviour of monetary authorities.

Key words: expectations hypothesis, monetary policy, changepoints

JEL classification numbers: E43

Rahapolitiikkaa koskevien pitkän aikavälin käsitysten mittaaminen ja korkojen aikarakenne

Suomen Pankin keskustelualoitteita 12/2004

Nicolas Rautureau
Tutkimusosasto

Tiivistelmä

Tällä tutkimuksella on kaksi tavoitetta. Ensimmäinen on mitata yleisön käsityksiä pitkällä aikavälillä harjoitettavasta rahapolitiikasta. Toinen tavoite on määrittää näiden käsitysten suhde pitkiin bondikorkoihin. Saksalaista aineistolla käytettäessä korkojen aikarakenteen mallintaminen kahden faktorin mallilla johtaa parhaisiin pitkien korkojen ennusteisiin tammikuusta 1975 tammikuuhun 2003 ja mahdollistaa myös pitkän aikavälin inflaatio-odotusten käyttämisen aikarakenteen toisena faktorina. Tämä faktori luonnehtii samalla vallitsevia käsityksiä rahaviranomaisten käyttäytymisestä.

Avainsanat: odotushypoteesi, rahapolitiikka, regiiminmuutokset

JEL-luokittelu: E43

Contents

1 Introduction

Long term interest rates play an important role in economics and finance due to their impact on real activity. Nevertheless, it is widely recognized that forecasting their level is difficult. In the same way, modelling their link to monetary policy is not so easy even if this relationship appears crucial, as emphasized by Goodfriend (1993) for the United States. Similarly the same difficulty concerns the rational expectations hypothesis framework in which most empirical analyses of the term structure of interest rates are conducted. This approach postulates that long term rates are a weighted average of current and expected future short rates plus a constant term premium. Even if numerous studies have presented evidence against the validity of the rational expectations theory in the past, from the start of the nineties, a growing number of positive results have been obtained and expectations about the short rate dynamics appear to be the main factor, but not the only one, driving the evolution of interest rates.

Moreover, the necessity of taking into account monetary policy in the rational expectations framework has been demonstrated in particular by the seminal work of Mankiw and Miron (1986) for the short end of the term structure. For the long part of the term structure, Fuhrer (1996), without specifying any process for the perception of regime shifts by agents, has shown that the expectations hypothesis can be accepted for the long part of the yield curve if we allow some small and discrete changes in the coefficients of the reaction function for the Federal Reserve System. These results confirm the relevance of monetary policy for the U.S. term structure analysis.

Four years after the launch of the euro in January 1999 and while the European Central Bank (ECB) plans a review about its monetary policy strategy, it seems interesting to study the link between monetary policy, the long-term perception of inflation and long-term interest rates. In the European case, it is possible to use the German term structure as a benchmark. But in this case, one feature is puzzling: how the long-term interest rates evolved in link with the expected path of the short-term interest rate, while this latter was under the control of the German monetary authorities until the end of 1998 (and normally depended of German macroeconomic conditions) and of the ECB after this date (and the euro area activity and inflation)? This question is important not only on a theoretical basis but also because two features are noteworthy for this period. The first is that we have observed a higher rate of inflation in Europe during the first years just after the launch of the euro in January 1999. The second feature is that recent studies on the subject conclude that euro interest rates are low relative to the levels derived from a reaction function for the Bundesbank. For example, Faust et al (2001) conduct a counterfactual exercise to compare the present policy of the ECB to the past one conducted by the Bundesbank. Hence from a reaction function with the Bundesbank parameters estimated with German series over the period 1985–1998 but with the euro area series they conclude that the ECB is more concerns on the output gap relatively to inflation, in comparison to the policy conducted by the Bundesbank in the past. Moreover, the estimates of Taylor rules by Gerdesmeier and Roffia (2003) from 1985 onward show some signs of structural changes in parameters around 1998. So it appears interesting to study the impact of the founding of the ECB on the public perception of monetary policy. In particular, the long-term perception, linked to the credibility of the policy, seems a particularly important topic.

This study contains two objectives. The first is the identification of the public perception of monetary policy to establish a relationship between this perception, the behaviour of monetary authorities and some key economic variables. The second objective is the identification of the relationship between monetary policy and the term structure of interest rates. In particular, we are interested by the link with long-term interest rates. From this perspective, the works of Kozicki and Tinsley (1998, 2001a, 2001b) are interesting for two reasons. Firstly, they show that conventional stationary and nonstationary specifications are unable to provide accurate forecasts of the short rate for long horizon and then are not suited to be used to fit long term interest rates, contrary to their own two-factor shifting endpoint representation. The idea behind their work is that forecasts could be dominated by intrinsic properties of the econometric specifications used. Secondly they propose a theoretical model to explain the existence of shifts in the stochastic process for the short rate and show how these shifts are linked to movements in the perception of monetary policy by the public. However, this demonstration is based only on the study of the stochastic process of inflation and on the hypothesis that, in the long run, inflation is a monetary phenomenon.

In the next section we outline some theoretical and empirical results about the relationships between monetary policy, the inflation target of monetary authorities, the level of this target perceived by the public and the long term interest rates dynamics. In section 3, we present the works of Kozicki and Tinsley (1998, 2001a, 2001b) and we establish the interest of this model in our framework. We also provide an illustration of the advantage of the shifting endpoint specification in comparison to traditional stationary and nonstationary specifications when autoregressive models are used to produce long-horizon expectations. In section 4 we propose two specifications for the long-term perception of the monetary authorities' inflation target to explain, with the Fisher relationship, the presence of shifts in the long-term perception of the short-term interest rate. Finally, some conclusions are drawn in section 5.

2 Monetary policy, inflation target and long term interest rates

Evans and Lewis (1995) and Crowder and Hoffman (1996) have shown that the Fisherian decomposition of a nominal rate into the expected real rate and expected inflation can be accepted if one includes in this relation a marginal tax rate τ on bond earnings. In this framework the Fisher relationship can be written:

$$(1-\tau)\,r_t = rr_t + \pi_t^e \tag{2.1}$$

where r_t, rr_t et π_t^e are the nominal interest rate, the real rate and the expected level of inflation respectively. Kozicki and Tinsley (2001a), Evans and Lewis (1995) and Crowder and Hoffman (1996) have proposed different estimates of the magnitude of the marginal tax rate for the United States. It appears that this rate can be set between 0.20% and 0.30%. Moreover the level of this tax seems to have fallen since the end of the eighties. But some authors confirm the one-for-one movement between inflation and interest rates without to take into account any marginal tax rate τ.

The relationship between inflation and interest rates has also been confirmed by numerous studies. We can summarise the results obtained in four points. The first is that in the long run, this relationship is nonlinear. For example, Evans and Lewis (1995) represent the inflation dynamics with a Markov process to take into account some episodic changes. More recently and in a similar way, Lanne (2002) finds, using a nonlinear bivariate mixture autoregressive model for the United States between 1952 and 2000, the presence of significant nonlinearities in the dynamics of nominal interest rate and inflation, while the real interest rate is stationary. Thus Lanne finds a one-for-one relationship between the nominal interest rate and inflation in the long run. For the United States and Canada, Tkacz (2002) uses two-regime threshold models to show how an asymmetric effect of monetary policy on inflation can explains the existence of a non-linear relationship between long-short yield spreads and inflation changes.

The second result underlines the interest of a two-factor model rather than a one factor model to represent the stochastic dynamics of interest rates. For example, Balduzzi, Das and Foresi (1998) show the presence of a stochastic central tendency in the dynamics of the short and long term interest rates which explains that the level of the short rate is not the only relevant variable to describe its conditional mean. For the United States between 1951 and 2001, King and Kurmann (2002) demonstrate the impact of this permanent component (stochastic trend) on the long-term interest rate dynamics. On the contrary, the spread depends more on the difference between the short rate and this stochastic trend, which is a temporary component.

The third result links this permanent component to monetary authorities behaviour. Recently, Gürkaynak, Sack and Swanson (2003) show that the change in expectations of the long-run inflation rate by private agents depend on macroeconomic and monetary surprises. Moreover, the relationship between inflation, interest rates and monetary policy has been studied for a long time and, for example, since the seminal paper of Mankiw and Miron (1986), a significant number of papers have studied the relationship between monetary policy and the rational expectations theory.

The last result concerns the existence of an asymmetric information set between private agents and monetary authorities so that the imperfect knowledge of the former explains why they adopt a learning strategy about monetary policy to infer the level of the long-term monetary policy target for inflation and why this process could take time and leads to some empirical facts like the excess sensitivity of long-term interest rates to the dynamics consistent with the rational expectations hypothesis. Hence Gürkaynak, Sack and Swanson (2003) show that since the mid-1997, when the Bank of

England gained more operational independence and the long-term inflation target was known, the dynamics of the long-term forward rate was more stable. Orphanides and Williams (2003) demonstrate also on a theoretical basis how the observed overreaction of long-term interest rates to the short-term interest rate could be explained by the presence of imperfect knowledge and a perpetual learning process by agents about the structure of the economy and the policymaker preferences. Hence, these two features increase the sensitivity of inflation expectations and of the term structure of interest rates to temporary economic shocks. As shown by Orphanides and Williams (2002), the imperfect knowledge is also responsible for the possible disconnection between the public's perception of inflation which result from the observed behaviour of monetary authorities, and the policy objective of the latter, while the same policy appears efficient under the rational expectations hypothesis.

The delayed response to economic shocks or to changes in monetary policy which characterize the inflation expectations and the interest rate dynamics is also explained by the existence of this asymmetric information set between private agents and monetary authorities. Hence, Huh and Lansing (2000) and Erceg and Levin (2001) are two models which demonstrate how an inflation scare problem could produce a sharp increase of long-term interest rates, as for the Volcker disinflation period of the early 1980s and the 1987 inflation scare episode. The sluggish decline of the inflation rate following these significant increases of long-term interest rates could then be explained by time necessary to private agents to perceived the shift to the new level of the long-term inflation targeted by monetary authorities. From a dynamic stochastic general-equilibrium (DSGE) model, Andolfatto, Hendry and Moran (2002) and Schorfheide (2003) demonstrate that the inflation and interest rates dynamics in the United States since 1960 could be characterized by a sluggish learning process about the policymaker's rule and the shocks that affect it.

3 Theoretical framework

At this level we need a model with two components and one constraint. The constraint is the ability to take into account the nonlinearities in the inflation and the term structure dynamics, a usual empirical fact in past studies and a possible outcome of the creation of the ECB. The first component is intended to identify the long-term perception of monetary policy by private agents. Here we use the long-term inflation target perceived by the public from the observed behaviour of the central bank and the Fisher relationship to link this perception to the corresponding level of nominal interest rate. The second component describes the relationship between the term structure of interest rates and this long-term perception of monetary policy. In this case the rational expectations hypothesis is the common transmission channel used. To take into account possible shifts due to some monetary policy regime changes, we use the shifting endpoint formulation of Kozicki and Tinsley (1998, 2001a, b). In the rest of this section we present the two-stage approach of Kozicki and Tinsley in a first stage. Then we explain our method to identify the second factor of the model, that is the public perception of the long-term inflation target. Finally, we present the rational expectations hypothesis as the transmission channel used in this study.

3.1 The two-stage approach of Kozicki and Tinsley

The approach of Kozicki and Tinsley is based on a two-step procedure. The first stage is to identify the endpoints, or the limiting conditional expectations of each variable of the model. The second stage is to include these endpoints in the autoregressive model for forecasting. The use of these endpoints serves to anchor the long term forecast of each variable of the model. Hence this approach introduces a second factor in the model and conducts to better forecasts than the traditional statistical models (stationary or nonstationary models).

This approach has several advantages in comparison to Structural VARs. First, this model doesn't use identification assumptions and the sensitivity of the results to them (see for example Evans and Marshall (1998) where three different empirical strategies are used). Second, it seems well suited in the European context around the founding of the ECB due to the common policy framework that could be defined by the same information set (ie a short term interest rate and two series for inflation and economic activity), the same goals of monetary policy but with possible shifts in the perception of monetary policy goals or in the level of these variables. As noted by Kozicki and Tinsley (2001b), these shifts are mapped into time-varying intercepts of the equation. Third, the definition of the endpoints is free, that is exogenous to the equation. For example, Kozicki and Tinsley compute the inflation endpoint from the analyse of the stochastic dynamics of inflation and the nominal short rate endpoint could be obtained directly from the term structure with a long-horizon forward rate or the infinite short rate of a zero-coupon yield curve. Finally, the study of the transmission process to the term structure is easy because on one side we have an estimate of an autoregressive model and, on the other side, we could see, with the expectations theory, the long term interest rates as an average of short term interest rates plus a constant term premium.

However, the difficulty of this approach is that this model has not been constructed in a first time to solve the identification problem of the public perception of monetary policy, since the definition of the endpoint is free, but to improve long-term forecasts with shifting perceptions of long term endpoints (or goals) of monetary policy. So, in respect to our two objectives, the definition of the endpoint should rest on a theoretical model which allows to identify the public perception of monetary policy[1].

[1] Otherwise we have to define additional assumptions. For example, in the VAR model of Kozicki and Tinsley the shifting endpoint for inflation is based on the stochastic process of the observed inflation, so they use the hypothesis that the inflation is under the control of monetary authorities to see this variable as the public perception of the long term inflation target of monetary authorities. But this hypothesis could eliminate the credibility problem.

3.2 Three specifications for the short rate process

Kozicki and Tinsley (1998, 2001a) deal with the univariate case to study the stochastic dynamic of the short rate. Hence, the general equation for the short rate r_t has the form

$$r_t = r_{t-1}^{(\infty)} + \sum_{i=1}^{m} a_i \left(r_{t-i} - r_{t-1}^{(\infty)} \right) + \varepsilon_t \tag{3.1}$$

where

$$r_t^{(\infty)} \equiv \lim_{k \to \infty} E_t r_{t+k} \tag{3.2}$$

This short rate endpoint $r_t^{(\infty)}$ could be seen as the limiting conditional central tendency of short rates. This variable could be subject to some shifts during the time with the evolution of the long-term public's perception of the short rate level.

The persistence of the short rate process (3.1) could be obtained by rewritten it in a mixed format of levels and differences:

$$\Delta r_t = \gamma \left(r_{t-1}^{(\infty)} - r_{t-1} \right) + \sum_{i=1}^{m-1} a_i^* \Delta r_{t-i} + \varepsilon_t \tag{3.3}$$

where $\Delta r_t = r_t - r_{t-1}$ and

$$\gamma = 1 - \textstyle\sum_{i=1}^{m-1} a_i$$
$$a_i^* = - \textstyle\sum_{j=i+1}^{m} a_j$$

γ can be interpreted as the speed at which the short rate reverts to its limiting conditional forecast (in the first term of (3.3), the short rate is incorporated in deviations from its long-run equilibrium value). Hence the persistence of the short rate increases when γ approaches zero and when this value is reached, we have a unit root process. Otherwise, a value of γ between (0,2) is consistent with mean-reversion. In fact, (3.3) has the form of the ADF test where the null hypothesis is *H0:* $\gamma = 0$ against $\gamma > 0$[2].

Three variations could be obtained from (3.3). The first is a constant endpoint specification (stationary representation). Here, $r_{t-1}^{(\infty)}$ is replaced by a constant $r^{(\infty)}$. This hypothesis corresponds with most theoretical models in finance, such as Cox et al (1985). In this case, the short rate follows a mean-reverting process toward the long-term value $r^{(\infty)}$. This latter is equal to the sample mean of the short rate in large sample[3].

[2]The ADF test has the form

$$\Delta r_t = c + \gamma r_{t-1} + \sum_{i=1}^{m-1} a_i^* \Delta r_{t-i} + \varepsilon_t$$

and concerns the hypothesis *H0:* $\gamma = 0$ against $\gamma < 0$.

[3]The long-term value $r^{(\infty)}$ is defined in this case by $r^{(\infty)} = c/\gamma$ in reference with parameters of the ADF test equation with a constant.

The second representation corresponds to a moving-average specification for the short rate, ie to nonstationary dynamics. In this case, the null hypothesis of a unit root process is accepted and the first term in (3.3) disappeared. This case is familiar in empirical macrofinance and has lead to numerous studies on nonstationary models and cointegration since the paper of Campbell and Shiller (1987, 1988).

The third specification is a shifting endpoint specification. In this case, the endpoint $r_{t-1}^{(\infty)}$ could vary with the time and the components of the underlying information set[4]. At this stage we have the choice between a fully specified information set to explain the movements of the short-rate endpoint or to limit the explanation power of the model by using only a forward rate directly observed from the yield curve (in this case we assume that the yield curve includes all the agents perceptions about the state of the economy and its future path). For example, in the latter case, $r_{t-1}^{(\infty)}$ could be obtained as the average of short rates at a long-horizon (for example between 5 and 10 years) or as the infinite short rate of a zero-coupon yield curve.

3.3 The definition of the nominal interest rate endpoint

As noted before, the endpoint definition in (3.3) is free. Hence, an *I(0)* assumption for x_t is consistent to a fixed endpoint, where its level is equal to the sample mean of the series. When we have an *I(1)* assumption, the endpoint is a moving average of past values observed. With a shifting endpoint specification, we can integrate some discrete shifts. Hence Kozicki and Tinsley obtain the endpoint for the short term interest rate from the implicit forward rate between the 5-year and 10-year observed yield for the United States:

$$\widehat{r}_t^{(\infty)} = \frac{D_{n'}(R_t^{n'} - \theta_{n'}) - D_n(R_t^n - \theta_n)}{D_{n'} - D_n} \tag{3.4}$$

where D_n and $D_{n'}$ are the duration for the 5-year and 10-year bonds and θ_n the constant term premium for the n-year bond[5]. For these maturities, the yield curve is relatively flat so (3.4) gives a good approximation of the long-term perception, at date t, of the short rate level. Another possibility to compute $\widehat{r}_t^{(\infty)}$ is to use the three-month forward rate in 10 years obtained from their zero-coupon curve, as in Jondeau and Sédillot (1999) for France and Germany. For other series the same approach can be applied. For example, for stationary series as the capacity utilization rate for manufacturing, the mean of the series is removed, implying a fixed endpoint of zero (see Kozicki and Tinsley (2001b)). For inflation, the endpoint process should include some episodic shifts due to

[4] Kozicki and Tinsley assume that the endpoint forecasts are fixed at each date t:

$$E_t r_{t+k}^{(\infty)} = r_t^{(\infty)}, \quad \forall k$$

[5] see for example Shiller, Campbell and Schoenholtz (1983). The duration is expressed for an i-period coupon-bearing bond by $D_i = \left(1 - g^i\right) / (1 - g)$, $i \geq 0$, where g is the discount factor. For an i-period zero-coupon bond, we have $D_i = i$.

changes in the agents perception of the inflation level, a common fact in the literature on this subject.

3.4 The transmission to the term structure

The transmission to the term structure of interest rates is obtained by the expectations theory. This hypothesis seems not to strong due to the recent positive results obtained and by the fact that even if the expectations hypothesis is not always statistically accepted, short rates expectations are identified as the main factor driven the shape of the yield curve. Hence, for zero-coupon bond yield, the long term interest rate $R_t^{(n)}$ is an average of the short term interest rates expected over the same period, r_{t+i}

$$R_t^{(n)} = \frac{1}{n} E_t \left[\sum_{i=0}^{n-1} r_{t+i} \right] + \theta_{n,t} \tag{3.5}$$

where $\theta_{n,t}$ is the term premium.

The use of the expectations theory means that monetary policy influences bond yields through a dependence of the expected path of short term interest rate on expected policy. In this framework, we use the short-rate equation (3.3) to generate short rate expectations in (3.5), from the information set available at date t. Hence, we use a weaker version of the expectations theory than the use of *ex post* values of the short rate. In this case, we have from (3.1) and with the notations of Kozicki and Tinsley (2001a)

$$E_t r_{t+i} = r_t^{(\infty)} + i_1' H^i \left(z_t - i_m r_t^{(\infty)} \right) \tag{3.6}$$

where i_m is an $m \times 1$ vector with all elements equal zero except for the first row which is equal to one, $z_t = [r_t, \ ..., \ r_{t-1-m}]'$ and

$$H \equiv \begin{bmatrix} a_1 & a_2 & \dots & a_m \\ & I_{m-1} & & 0_{m-1,1} \end{bmatrix}$$

Hence, if we substitute (3.6) in (3.5), we have

$$R_t^{(n)} = r_t^{(\infty)} + \left(\frac{1}{n} \sum_{i=0}^{n-1} i_1' H^i \right) \left(z_t - i_m r_t^{(\infty)} \right) + \theta_{n,t} \tag{3.7}$$

Equation (3.7) is a description of the contemporaneous relationship between bond yield and states variables. Predictions of yields consistent with the information set available in *t-1* can be written as

$$E_{t-1} R_t^{(n)} = r_{t-1}^{(\infty)} + \left(\frac{1}{n} \sum_{i=0}^{n-1} i_1' H^{i+1} \right) \left(z_{t-1} - i_m r_{t-1}^{(\infty)} \right) + E_{t-1} \theta_{n,t} \tag{3.8}$$

From (3.8) we can see that the importance of the endpoint definition increases with the forecast horizon (with the increase of i).

3.5 The forecast of the German long-term interest rates

We use the German series between 1975 and 1998 for the one-month short rate, and the euro area series from 1999 onward. The 5 and 10-year bonds are zero-coupon yields obtained from the Bundesbank over all the period. We choose the German series due to the central place of the German policy in the European Monetary System before the launch of the ECB and because the German bond rates remain the benchmark for the long-term interest rates in the euro area. The selection of the short-term interest rates then follows our choice for long-term interest rates and the expectations theory. Until 1998, the German bond rates evolved in link with the expected path of the German short-term interest rate and since 1999, their evolution depend on the expectations about the future behaviour of the euro-area short-term interest rate fixed by the ECB.

We start by showing on figure 1 the evolution of the one-month interest rate and of the 5 and 10-year bond rates. From 1975 until January 2003, we can see that the three series follow a similar pattern even if some differences exist, as between 1993 and 1999 where the decrease is more pronounced for the one-month interest rate than for bond rates. Figure 2 illustrates three measures of volatility for German long-term interest rates. The top panel represents the year-to-year change of the 5 and 10-year interest rates and show a sharp increase of the bond rates just after the launch of the ECB. Nevertheless, this increase does not appear exceptional in its magnitude in comparison to previous periods. The middle panel shows the volatility of the bond rate over 12 months. This is computed at the date t as the standard error of the series over the interval $[t\text{-}12,\ t]$. And the bottom panel shows the same measure but where the series is the long-term minus the one-month interest rates to eliminate any tendency and level impact. In the two cases, the volatility appears relatively high during the first years after the launch of the ECB in comparison to the nineties, particularly for the third measure but without any excess, especially if we compare to the results obtained since 1975[6]. Hence this instability over nearly 30 years offer an interesting framework to analyze the impact of the short-rate process specification in the forecast of German long-term interest rates.

We compare now the performance of the three specifications for the short rate process in the forecasting of long-term interest rates. We start by computing the nominal interest rate shifting endpoint directly from the yield curve from (3.4). We proceed in two stages as in Kozicki and Tinsley (2001a). First we obtain a first approximation of the endpoint $\widehat{r}_t^{(\infty)}$ by ignoring the different constant term premia with the hypothesis that $D_{n'}\theta_{n'} - D_n\theta_n$ equal to zero. Second, we proceed to a constant adjustment to equalize the mean of the shifting endpoint estimate to the one of the short-term interest rate on the sample. Result is shown on figure 3 under the label 'shifting endpoint' and significant fluctuations of $\widehat{r}_t^{(\infty)}$ appear between 1975 and 2003.

We estimate after that the short-rate equation (3.3) with 6 lags and where a constant a_0 is added for the estimate. The three endpoint specifications

[6] These measures are also influenced by business cycles.

based only on the term structure appear in the column 2–4 in table 1. For the stationary endpoint specification $r_{t-1}^{(\infty)}$ is replaced by a constant $r^{(\infty)}$ which has been also estimated. For the moving average specification, the parameter γ is restricted to equal zero. $\widehat{r}_t^{(\infty)}$ is computed as describe in the previous section for the shifting endpoint specification.

Table 1: Autotoregressive models of the short rate. 1975–2003

	Endpoint characterization				
	Constant	Moving average	Shifting endpoint	Learning model	Kalman filter model
a_0	0.106 (0.054)	−0.003 (0.021)	−0.0007 (0.021)	−0.026 (0.021)	−0.012 (0.017)
γ	0.020 (0.009)		0.030 (0.010)	0.028	0.062 (0.040)
$\sum_{i=1}^{m-1} a_i^*$	0.407	0.355	0.434	0.691	0.958
R^2	0.057	0.047	0.069	0.089	0.092
$RMSE$	0.376	0.378	0.373	0.155	0.155
	RMSE of the long-term bond rate prediction				
5-year	1.586	1.874	1.265 / 0.969	0.967	1.017
10-year	2.014	2.372	1.565 / 1.128	1.120	1.094

For the constant, moving average and the first number of the shifting endpoint specifications, the RMSE of the long-term bond rate prediction concerns the period 1975:01–2003:01. For the second number of the shifting endpoint specification, the learning and kalman filter models, this statistic is computed over the period 1995:01–2003:01.

Two comments could be made for the constant endpoint specification. First, the value of γ falls in the range (0, 2) so this result confirms that the short rate follow a mean-reverting process. Second, the implied short rate endpoint over the sample is 5.42% (a_0/γ). The nonstationary representation is estimated under the constraint $\gamma = 0$. The Root Mean Squared Error (RMSE) statistic is closed to the level observed for the constant endpoint specification which means that a unit root seems to be present in the short rate process. This is confirmed with the ADF-test which has a similar form as the constant endpoint specification of table 1. Hence, the t-statistic for γ in this specification, with a value of 2.22 does no reject the null hypothesis that $\gamma = 0$ (the critical value is 2.56 at a 10%-level). Results for the shifting endpoint specification confirms the mean-reverting process followed by the short rate. And the highest value of γ shows an increase in the convergence speed to the endpoint. Finally, the closed values observed for the R^2 and the RMSE statistics for the three models show that the results for a one-period ahead prediction of the short rate do not depend significantly on the specification of the endpoint.

From these results, we compute the one-period-ahead long-term interest rate forecasts using the rational expectations theory and under the implicit hypothesis that an autoregressive process like (3.3) is well suited to model the expectations of the future path of the short-term interest rate. Predictions of the expectation component in the long-term yields consistent with the information set available in *t-1* are obtained from (3.8) where the term premium θ is fixed to zero. The interest of the shifting endpoint representation to forecast the 5-year and 10-year bond rates is clear on figures 4 and 5. While the forecast of the 5 and 10-year bond appears too stable with the stationary representation and too volatile with the nonstationary representation, the result appears better with the shifting endpoint specification. This is confirmed by the last two rows of table 1 with the RMSE statistic (the first number for the shifting endpoint specification). Hence, the difference between the actual and the fitted series represents the residual or the constant term premium (under the assumption that there is no error in the statistical representation of the short-rate expectations). For the shifting endpoint specification, the average value of this premium over the sample is 1.112% for the 5-year bond rate and 1.519% for the 10-year bond rate[7]. In summary the shifting endpoint specification introduces a second factor in the dynamics of interest rates and this allows to take into account the observed nonlinearities in their observed behaviour.

4 Two approaches to link the shifting endpoint to the inflation target perception

The issue of interest in this section is to provide a theoretical basis for the existence of a shifting endpoint in the short rate forecasting model because, if the advantage to use the shifting endpoint specification appears clearly in the previous section, the use of an implicit forward rate does not help to identify the link between this long-term perception of the future level of the short rate and some particular economic information. A natural explanation comes from the Fisher relationship which link the nominal interest rate to an expected real component and an expected inflation rate. Evans and Lewis (1995), Crowder and Hoffman (1996), Kozicki and Tinsley (2001a) and Lanne (2002) obtained two results. First, when anticipated shifts in the inflation process (and sometimes a marginal tax rate) are taken into account, there exist a one-for-one long-run movements between nominal interest rates and expected inflation in the long-run. Second, innovations in the inflation process are the likely source of nonstationarity of nominal interest rates and the real rate is not a significant source of shifts in nominal interest rates. Moreover, Paloviita (2002) finds that inflation expectations are the main factor leading the inflation process in all euro area countries.

In this framework, it is possible to link movements in the the nominal rate endpoint to the long-term level of inflation and indirectly to the monetary

[7]Kozicki and Tinsley (2001a) estimate this constant term premia under the hypothesis that the short rate process is homoskedastic (equation (15) in their paper). They report the expected value of the long-term nominal interest rates (which include this constant term premium). They obtain a value very close to our results for the magnitude of these constant term premia (1.264% and 1.338% for the 5 and 10-year bonds respectively).

authorities' perceived inflation target by the public. So either we consider that it is enough to study only the inflation process to obtain an estimate of the inflation endpoint or this latter should be linked to an explicit model of monetary authorities behaviour. In the second case, that allows the inflation process to be out of the control of monetary authorities, or in a less extent, the imperfect knowledge of private agents leaves open the possibility to observe a disconnection between the public's inflation perception which result from the observed behaviour of monetary authorities, and the policy objective of the latter. The end of this section rests on two stages, where each time, the Fisher relationship is used to obtain the corresponding nominal interest rate endpoint. First the inflation endpoint is obtained directly from the observed stochastic dynamics of inflation. Second, shifts in the inflation endpoint process come from the observed behaviour of the monetary authorities to introduce explicitly two common features of the literature on the subject, that is the role of monetary policy and the existence of an asymmetric information set between private agents and central banks.

4.1 Long-run inflation expectations and the stochastic dynamics of inflation

The previous section has documented first the interest of the shifting endpoint specification in comparison to traditional autoregressive specification and second, the presence of some shifts in the short-term nominal endpoint stochastic process, or in other words, in the long-term perception of the level of the short-term interest rates. In the appendix of this paper, we have also applied the break point test methodology developed by Bai and Perron (1998a, b) to see if these shifts in the long-term perception of the short-term interest rate level go with some shifts in the inflation process. The sequential estimate of an unknown number of breaks points conclude to the presence of such changes in the dynamics of inflation. Hence, with the Fisher relationship, it could be possible to explain the shifts in the short-term nominal endpoint stochastic process by changes in the long-run inflation expectations.

The issue of interest in this section is the public perception of the shifts occurred in the inflation stochastic process in Europe since 1975 and its modelling. This procedure has to allow the obtaining of an inflation endpoint to built next a nominal interest rate endpoint and to forecast German long-term interest rates. To answer to this question, we follow the works of Kozicki and Tinsley (2001a, b) and we apply a 'real-time' process to detect structural changes in inflation dynamics, with different degrees of monitoring to introduce heterogeneity between agents. This method rests on the statistical test developed by Andrews (1993) and Andrews and Ploberger (1994) and allows to take into account the different common features of the literature, namely an asymmetric information set between private agents and monetary authorities and the presence of episodic changes in the inflation process. Moreover, as emphasize by Gerberding (2001) for example, these phenomena take into account the 'stickiness' of the inflation process by introducing a delay of time before to check for a new break in the inflation dynamics. The inflation series from 1975 onward is built from the CPI series for Germany before 1999 and from the HICP series for the euro area after this date.

We use in this section the regression based approach of Kozicki and Tinsley (2001a). The model rests on three hypothesis[8]. First information is asymmetric between agents and monetary authorities so that private agents must watch the behaviour of the central bank, or its apparent consequences, to detect an episodic shift in the inflation long-term target of monetary authorities[9]. Second learning occurs in real time but agents need time to collect and analyze information, so there exist a recognition lag between the period where the shift occurs and the period where this shift is taken into account by private agents. Third, the economy is composed of heterogeneous agents, more or less active to monitor the behaviour of monetary authorities. This degree of activity is modelled by playing with the number of periods an agent accumulate from the last detected shift before to test again for the presence of a new structural break in the inflation process[10].

Between two structural breaks, the endpoint of the inflation process is constant. Without any shift in the long-term perception of inflation over the period, the constant endpoint $\pi^{(\infty)}$ could be obtained from the following equation in link with (3.3)

$$\Delta\pi_t = c - \gamma\pi_{t-1} + \sum_{i=1}^{m-1} a_i^* \Delta\pi_{t-i} + \varepsilon_t \qquad (4.1)$$

where the constant endpoint $\pi^{(\infty)}$ is equal to $\pi^{(\infty)} = c/\gamma$.

But to take into account the presence of episodic shifts in the endpoint level over all the period, the stochastic dynamics of the inflation process π is studied from the following shifting endpoint specification where the structural changes are taken into account through a change in the constant level c. In this case the autoregressive model for inflation is

$$\Delta\pi_t = c + \sum_k \alpha_k \delta\left(t-k\right) - \gamma_\pi \pi_{t-1} + C(L)\Delta\pi_{t-1} + a_t \qquad (4.2)$$

where $C(L)$ is a polynomial in the lag operator and k=$[k_1, k_{2,\ldots}]$ indexes periods of change in the inflation target of monetary authorities. When a new change is detected in the constant, the binary dummy variable $\delta\left(x\right)$ switches from zero to one. As in the stationary autoregressive case, the policy endpoint in

[8]We present only the model we use in this section. In their other paper, the aggregate endpoint series is obtained from a stochastic discrete choice model.

[9]The difference between this section and the approach of the next section rests on this point, with a model limited to the monitoring of the inflation dynamics in one case and a model where the policy target is infered from an explicit model for monetary authorities behaviour in the other case.

[10]This number of periods before to control a new time the behaviour of monetary authorities through the control of inflation dynamics could be linked to the trimming parameter ϵ in the previous section.

period τ is defined by

$$\widehat{\pi}_{\tau}^{(\infty)} = \left[c + \sum_{k<\tau} \alpha_k\right] / \gamma_{\pi} \tag{4.3}$$

but in this case, episodic structural changes in the constant are allowed which have an impact on the estimated long-term level of inflation $\widehat{\pi}_t^{(\infty)}$.

The detection of changes in the long-run policy target is obtained by implementing an expanding-version of sequential searches for multiple changepoints. Hence, agents start with an initial sample. They wait a fixed number of periods (different between heterogeneous agents) to apply a maximum-Wald test procedure to detect if there is a shift in the endpoint of (4.2), as in Andrews (1993) and Andrews and Ploberger (1994)[11]. This latter statistic has been employed for example by Benati and Kapetanios (2002) to test for the presence of multiple structural breaks at unknown points for the inflation dynamics of 18 countries and the euro zone. To detect a shift in the endpoint of (4.2), agents start with an initial sample. Then they add a new observation to this sample and, after the exclusion of the 5% of observations at the two extremities of the sample, they test for a change in the constant parameter in (4.2), the other parameters remaining constant. We then focus on the *ave*-Wald version of the Andrews (1993) and Andrews and Ploberger (1994) test[12]:

$$ave{-}Wald = \frac{1}{(T_2 - T_1 + 1)} \sum_{t=T_1}^{T_2} Wald(t) \tag{4.4}$$

We select six newsletters which define six recognition lags from one to six years (with a regular 12-month interval between each newsletter) and this approach is applied for each of them. Figure 8 represents two examples of these newsletters. With the increase of the recognition lag, the delay between a change in the inflation process and a change in the perceived long-term level of inflation increases also. But no significant change is perceived since 1996 in the policy target.

The aggregation of the different breakpoints allowed by the model is obtained from a regression approach which gives the distribution of each categories of agents in the endpoint perception of the nominal interest rate

$$\widehat{r}_t^{(\infty)} = \alpha + \sum_{i=1}^{6} w_i \widehat{\pi}_{i,t}^{(\infty)} + \varepsilon_t \tag{4.5}$$

where $\widehat{r}_t^{(\infty)}$ is the shifting endpoint estimate obtained directly from the term structure with (3.4), $\widehat{\pi}_i^{(\infty)}$ is the endpoint estimated for the newsletter i and w_i are the weights of each letter i, i=1, ..., 6. The weights lie on a third-degree

[11] The 5% critical values are computed by these authors.

[12] We concentrate on the *ave*-Wald test version of Andrews (1993) and Andrews and Ploberger (1994) because the Wald version of this test seems to exhibit more power than the likelihood ratio or Lagrange multiplier versions (see Benati and Kapetanios (2002)).

polynomial to smooth their distribution and avoid negative values. This approach conduct to the following result

$$\begin{aligned}\widehat{r}_t^{(\infty)} &= 2.033 + 0.145 * \widehat{\pi}_{1,t}^{(\infty)} + 0.107 * \widehat{\pi}_{2,t}^{(\infty)} + 0.019 * \widehat{\pi}_{3,t}^{(\infty)} \\ &\quad + 0.00001 * \widehat{\pi}_{4,t}^{(\infty)} + 0.169 * \widehat{\pi}_{5,t}^{(\infty)} + 0.646 * \widehat{\pi}_{6,t}^{(\infty)} \end{aligned} \quad (4.6)$$

The nominal endpoint is then obtained as the fitted value of (4.6) and is represented on the top panel of figure 3 under the label 'Learning endpoint'. The overall dynamics of the learning endpoint is closed to the one of the shifting endpoint computed from the yield curve, even if a time-lag exists. From (4.6) the real interest rate estimate is the constant, that is 2.033% if we exclude any marginal tax rate τ in (2.1). In the other case, the level of the tax is defined as $\alpha/\sum_{i=1}^{6}\widehat{w}_i$. The sum of the estimated weights equals 1.086 and this corresponds to a tax level $\widehat{\tau}$ of 0.079% and to a real interest rate equals to 1.872%[13]. The mean detection lag, computed as the product of the estimated weights by the different recognition lags conducts to an estimate of more than 5 years (5 years one month and 19 days). This lag could explain why the learning endpoint does not increase since 1997 despite an increase in actual inflation[14].

The forecast of the five and ten-year German bond rates from the Learning endpoint is then obtained first from the estimate of the short-term interest rate process (3.3) where $r_{t-1}^{(\infty)}$ is replaced by the Learning endpoint estimate and result appear in the fifth column of table 1. Parameters are very closed to the shifting endpoint estimate. Then, the five and ten-year German bond rate predictions are obtained as in section 3.5 from (3.8) where the term premia θ is fixed to zero. Result of this procedure is reported on the middle panel of figure 11 since 1995. The forecast is very closed to the shifting endpoint solution. This is confirmed by the RMSE statistic in the last row of table 1, where the second value for the shifting endpoint specification is similar to the Learning endpoint specification.

4.2 Long-run inflation expectations and the observed behaviour of monetary authorities

In this section we take into account explicitly another distinctive feature of the literature, namely the monetary policy influence on the long-term perception of inflation and on the dynamics of long-term interest rates. Hence the approach does not rest only on the inflation dynamics but on the inflation level consistent with the observed behaviour of the central bank, through the observed dynamic of the short-term interest rate. The choice of the short-term interest rate as the instrument of central banks is a common feature of the literature and is closely linked to the interest rate setting described by Blinder (1998) for example:

[13] This value seems slightly lower to the real interest rate on the period. For example, the difference in the sample between the one-month interest rate and the monthly inflation on an annual basis is 2.58%

[14] Kozicki and Tinsley (2001a) find also a mean recognition lag that exceed five years for the US case.

> “The central bank must adjust its nominal interest rate so as to guide the real rate back toward its neutral setting. (...) I propose to define the neutral interest rate as the interest rate that equates GDP along this steady-state IS curve to potential GDP. (...) Thus the proposed definition of neutrality is oriented entirely toward the control of inflation, as seems appropriate given that price stability is the primary long-run responsibility of any central bank. (...) My suggestion, then, is that central banks estimate the neutral interest rate on a regular basis (a range may make more sense than a point estimate) and use that estimates as the ‘zero point’ on their monetary policy scales. Any higher interest rate constitutes ‘tight money’; any lower rate constitutes ‘easy money’. Neutrality is the only viable policy setting for the long run”[15].

This method allows the estimate of the implicit level of inflation and the examination of its link with the observed inflation and with the long-term interest rates dynamics. In this framework, for a credible central bank and with a symmetric information hypothesis, the long-run equilibrium value for inflation can be considered as the long-term inflation objective pursued by monetary authorities. But some differences could exist between the perceived and the official target level of inflation if these conditions are not fulfilled. In this section, the framework allows to relax the hypothesis that inflation is under the control of monetary authorities. Hence, if the results are closed to those obtained from the examination of the inflation series alone, this will translate a good credibility of the central bank, ie that the actual inflation process does not differ to the dynamics implied by the central bank behaviour.

4.2.1 Forward looking reaction function and rolling regressions

In this section, private agents observe the behaviour of monetary authorities through the rule followed by the central bank, where a policy rule is defined as a sequence of short term interest rates. Bomfim and Brayton (1998) use a similar framework to link the long-term perception of the policy target to the policy rule. Their approach is modified by taking into account a forward looking specification for the policy rule, in a similar fashion than in Clarida, Gali and Gertler (1998). Hence the underlying theoretical framework links the short-term interest rate i_t to two elements: a target interest rate i_t^* and a smoothing parameter ρ to limit the fluctuations of i_t

$$i_t = (1-\rho)i_t^* + \rho i_{t-1} + \varepsilon_t \tag{4.7}$$

The target interest rate satisfies the equation

$$i_t^* = \alpha + \beta E_t(\pi_{t+n}) + \gamma E_t(y_t) \tag{4.8}$$

[15] Blinder (1998), p. 33–35.

where π_t is inflation, y_t an activity gap measure and n the forward looking horizon. With (4.8) in (4.7) we have

$$i_t = (1-\rho)\alpha + (1-\rho)\beta\pi_{t+n} + (1-\rho)\gamma y_t + \rho i_{t-1} + \upsilon_t \quad (4.9)$$

Estimate of (4.9) is then conducted from

$$i_t = \theta_1 + \theta_2\ \pi_{t+n} + \theta_3\ y_t + \theta_4\ i_{t-1} + \upsilon_t \quad (4.10)$$

where the forward looking horizon is fixed to 12 months.

In the long-run one can assume that the real rate of interest (rr^*) is known and constant, that the equilibrium nominal rate of interest moves one-for-one with equilibrium inflation and that the unemployment gap is zero. It is further assumed that nominal interest rates like inflation rates are constant over time ($r_t = r_{t+i} = \overline{r}$ and $\pi_t = \pi_{t+i} = \overline{\pi}$, $\forall i = 1, ..., \infty$). In this case, the implicit target rate consistent with (4.10) is given by

$$\pi^{(\infty)} = \frac{\theta_1 - (1-\theta_4)rr^*}{1 - \theta_4 - \theta_2} \quad (4.11)$$

The corresponding endpoint perception for the short-term nominal interest rate $r_{t,rf}^{(\infty)}$ could then be obtained from the addition of $\hat{\pi}^{(\infty)}$ with an hypothesis for the level of rr^*. Estimate of (4.10) are conducted from a rolling regressions approach with a constant ten-year window. GMM and two stages least squares are used with six lags for each of the three exogenous variables in (4.10).

The interest rate series i_t is the one-month interest rate for Germany between 1985 and 1998 and the one-month interest rate for the euro area afterwards, as we also want to explain the behaviour of the German interest rate. The inflation and activity series are built in the same fashion. We use mixed data, that is German data until 1998 and euro area data after that[16]. This procedure implies another choice for the rescaling of the data. We do not rescale inflation series because the short-term interest rate depends on the actual level of inflation[17]. The CPI series is employed for Germany before 1999 and is combined either with the HICP series or with a core measure of inflation since 1999. Activity gap measures are built from the industrial production index with an Hodrick-Prescott filter (*OgapHp*) and with a quadratic trend (*OgapTrd*) for the Output gap. The unemployment gap series (*Ugap*) comes from the OECD. All these series are monthly data except for the quarterly frequency of the unemployment gap, for which we have supposed a constant

[16] The question concerns the relative importance of national and euro area macroeconomic series to determine national long term interest rates. We do not aggregate data for the euro area from 1985 onward (with an artificial euroland before 1999), as in Gerdesmeier and Roffia (2003) for example because we want to explain the dynamics of long-term German bond rates and these latters were dependent on German macroeconomic conditions before 1999.

[17] This problem is limited by the convergence period across euro area before the launch of the ECB and by the fact that German bond rates are the benchmark for European long-term interest rates. Moreover if a significant change had occured in the level of inflation after the launch of the ECB, this would have led an increase in the short-term interest rate level. We adopt here the position of a German investor monitoring the behaviour of monetary authorities to forecast long-term German bond rates.

value on the quarter. Series with and without rescaling have been employed but the difference are minor due to limited value of θ_3 in (4.10) and the importance of the smoothing parameter[18]. All these series are represented on figure 9.

Different specifications have been tested concerning among others the length of the window, the number of lags or the real rate level hypothesis rr^* and results are reported in the first eight rows of table 2. For all these results, the ten-year window used means that inflation target perception is available since 1995. The Root Mean Square Error statistics is computed between $\widehat{r}_t^{(\infty)}$, the shifting endpoint specification obtained directly from the term structure and $r_{t,rf}^{(\infty)}$, the short-term interest rate endpoint perception implied by the rolling regressions[19]. Figure 10 reports also some estimates of $r_{t,rf}^{(\infty)}$. Two conclusions could be drawn. First this method is not robust to the choice of the series and to the estimation method, which explain non plausible values for $\widehat{r}_{t,rf}^{(\infty)}$ observed in some cases (top panel of figure 10) and make difficult any robust inference on the long-term inflation expectations. Second some estimates conduct to better result, as it is the case when core measure of inflation and the HP filter method to build output gap are used (middle panel of figure 10).

Table 2: RMSE for the long-run short interest rate measures

Name	Real rate	Series		Method	RMSE
		Inflation	Activity		
Frwrd1	2%	HICP	OgapHP	GMM	2.376
Frwrd2	2,5%	HICP	OgapHP	GMM	2.920
Frwrd3	2%	HICP	OgapTrd	GMM	37.260
Frwrd4	2%	HICP	Ugap	GMM	2.419
Frwrd5	2%	core	OgapHP	GMM	2.088
Frwrd6	2,5%	core	OgapHP	GMM	1.975
Frwrd7	2%	core	Ugap	GMM	4.617
2SLS	2,5%	core	OgapHP	2SLS	1.894
Kalm1	2%	HICP	OgapHP	Kalman	1.257
Kalm2	2%	core	OgapTrd	Kalman	1.257

4.2.2 Kalman filter

To get round these difficulties to estimate this unobservable variable which is the perceived inflation target followed by the European monetary authorities, we use the Kalman filter approach. This methodology has been used for example by Laubach, Williams (2001) for the natural rate of interest, potential output and its trend growth rate in the United States and by Lenz (2001) for

[18] The value of $\widehat{\rho}$ in the rolling regressions is around 0.90 and less than 0.05 for $\widehat{\theta}_4$.

[19] The two measures could not be too different if we have consistent estimates.

the monetary policy stance in Switzerland. In this section we use the same approach than Bomfim and Brayton (1998)

$$i_t = \beta_1 i_{t-1} + \beta_2 (\pi_t - \pi_t^*) + \beta_3 y_t + \beta_4 y_{t-1} + \beta_5 (rr^* + \pi_t) + e_t \tag{4.12}$$

with

$$\pi_t^* = \pi_{t-1}^* + \varepsilon_t \tag{4.13}$$

where $\beta_1 + \beta_5 = 1$. e_t and ε_t are uncorrelated and follow an i.i.d. process. The state-space representation of this model takes the form

$$i_t = x_t' \Gamma_t + e_t \tag{4.14}$$

with

$$\Gamma_t = \Gamma_{t-1} + \eta_t \tag{4.15}$$

where $\mathrm{x}_t \equiv [1,\ i_{t-1},\ \pi_t,\ y_t,\ y_{t-1}]'$ and $\Gamma_t = [\beta_5 rr^* - \beta_2 \pi_t^*,\ 1 - \beta_5,\ \beta_2 + \beta_5,\ \beta_3,\ \beta_4]'$. η_t is a vector of zeroes, except for its first element so only the inflation target varies stochastically.

The long-term inflation perception in this case is assumed to change when a difference higher than 0.25 point is detected from the previous period for π_t^* in (4.13). The result of two specifications is reported in the last two rows of table 2 where the short interest rate endpoint is obtained by the addition of the real interest rate to the long-term inflation perception. But due to the importance of inertia in (4.12) with a value of β_1 closed to 0.90, modifications concerning π_t and y_t have no significative impact.

The analysis based on RMSE shows than the Kalman filter methodology allows a better approximation of the endpoint obtained from the term structure than with rolling regressions. The reason is that this method does not conduct to observe extreme values. The endpoint series obtained with the Kalman filter and reported on the bottom panels of figure 10 reveal the presence of a learning period, which means that expectations need time to adjust to the shifts observed in the term structure. We can also see than the perceived inflation target of the ECB, which is obtained by subtracting the real interest rate level of 2%, exceeds the 2% level between 2001 and the start of 2002 (ie the nominal short rate endpoint exceeds 4%). But the end of the sample coincides with a return under the 2% level.

Figure 3 reports the three different nominal endpoints obtained with the Fisher. We recall that the 'shifting endpoint' label concerns the endpoint obtained directly from the term structure in section 3.5, the 'Learning endpoint' model has been estimated only from the observed inflation dynamics with equation (4.6) and the 'Kalman filter' label concerns the endpoint obtained in the previous section. We can see a close correspondence between these three series, especially for the tendency even if some significant differences appear punctually. Figure 11 then shows the forecast of the 5 and 10-year German bond rate with this three endpoint measures include in (3.3) and (3.8). The RMSE statistic of the 5 and 10-year German bond rates forecast for the period 1995:01 to 2002-12 in the last row of table 1 confirms the good ability of this specification to anticipate the level of the long-term interest rate.

5 Conclusion

This study pursued two objectives. The first was the identification of the public perception of monetary policy to establish a relationship between this perception, the behaviour of monetary authorities and some key economic variables. The second objective was the identification of the relationship between monetary policy and long-term interest rates. At the end, four main results appear. First, the shifting endpoint specification seems well suitable to forecast German long-term interest rates. This result is not a surprise and confirms those obtained by Kozicki and Tinsley (1998, 2001a, b) and Jondeau and Sédillot (1999). With these two factors models the risk premia appear more stable over time and the average value is estimated between 1.1 and 1.5% for the 5 and 10-year bond rate.

Second, the use of the Bai and Perron (1998a, b) methodology to detect structural changes in the inflation stochastic process also confirmed the impact of shifts in the long-term inflation perception on the short-term interest rate dynamics and on its long-term perception. From this point of view, the increase in the level of inflation between 1999 and 2002 appears significant while remaining reasonable in an historical perspective.

Third, the short-term interest rate endpoint could be linked to the long-term perception of inflation and to the monetary authorities policy. We obtain this result in two ways, from a model with heterogeneous agents controlling the stochastic process of inflation and from a model where the implicit long-term perception of inflation is built by inverting a reaction function of monetary authorities. In the two cases, the Fisher relationship is used to obtain the corresponding nominal interest rate endpoint. Moreover, the forecast of German bond rates since 1995 using these short-term interest rate endpoints as second factor conduct to similar results than those obtained when this second factor is computed directly from the term structure with the shifting endpoint specification. This confirms the validity of the two approaches followed here and the importance of inflation expectations and monetary policy as factors driving the long-term interest rates dynamics.

Finally, from the Kalman filter specification, the inflation expectations return under a 2% level from the middle of 2002, what seems to be an indication of the credibility of the ECB. But this point requires more advanced works to be validated.

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Appendix

An ex post approach to test structural changes in the inflation dynamics

The endpoint of a stationary model is constant at a level equals to its sample mean in large samples. Hence in this simple model, we are interested to detect structural changes in the mean of the inflation series. To do that, we proceed as in Bai and Perron (1998b) where they study the U.S. ex-post real interest rate dynamics. Hence, we suppose that agents detect changes in the mean of the inflation series and consider that until a new structural change is detected, the best forecast for the long-term perception of inflation is the observed mean of the present 'period', where a 'period' is defined as the time elapsed between two structural changes in the inflation process. Bai and Perron (1998a, b) consider the general multiple linear regression framework with m breaks

$$\begin{array}{ll} \pi_t = x_t'\beta + z_t'\delta_1 + u_t, & t = 1, \ ..., \ T_1 \\ \pi_t = x_t'\beta + z_t'\delta_2 + u_t, & t = T_1 + 1, \ ..., \ T_2 \\ \vdots & \\ \pi_t = x_t'\beta + z_t'\delta_{m+1} + u_t, & t = T_m + 1, \ ..., \ T \end{array} \tag{A1}$$

where π_t is the observed level of inflation at time t and T the number of observations. x_t and z_t are vectors of covariates with $(p\times 1)$ and $(q\times 1)$ elements respectively and β and δ_i are the corresponding vectors of coefficients. u_t is the disturbance and the indices $(T_1, T_2, \ldots, T_m)$, that is the break point dates are unknown. We observe a pure structural change model only when p=0. Hence Bai and Perron (1998b) propose a method to estimate together the unknown regression parameters β and δ_i with the break points $(T_1, T_2, \ldots, T_m)$. Our test about the stability of the inflation process mean conducts to choose $x_t = 0$ and $z_t = \{1\}$ in (A1).

The first outcome of this procedure concerns the number of structural breaks observed in the sample. This is done with the sequential application of a supF_T $(l+1|l)$ test where the null hypothesis is the absence of a new structural break. The procedure starts with a zero-break solution. Then tests for the presence of different subsamples are conducted from parameters constancy tests. When a rejection of a supF_T $(l+1|l)$ test is observed for a subsample, a break is added. The number of breaks l increases sequentially until the test fails to reject the null hypothesis. For a model with l breaks, the date of the estimated structural breaks $(\widehat{T}_1, \ldots, \widehat{T}_m)$ are obtained by a global minimization of the sum of squared residuals. The presence of another structural change is accepted if the latter sum for the $(l+1)$ breaks solution is significantly smaller than the sum for the l breaks solution. The critical values are computed by Bai and Perron (1998a, b).

This method is an ex post method to detect break points and could not be used in real time. But this gives a useful indication to detect break points for inflation in the European context. The result depends also on a trimming parameter ϵ, that is, the percentage of the sample which defines the minimum sample length. The maximum number of breaks allowed, which is not in all cases the maximum number of breaks observed, is bounded to 9. We model the heterogeneity between private agents by the attention they bear to monitoring the dynamic of inflation[20]. Hence an active agent is subject to important costs due to inflation and accumulate less data before to test again for the presence of a new break. The five possibilities for the trimming parameter ϵ are $\{.05, .10, .15, .20, .25\}$. They define five categories of agents and as many sensitivities[21]. For example, we have 335 data in our sample so the minimum number of periods accumulated before to test for a new break varies between 16 and 83 months ($Integer(\epsilon \times T)$).

The results of the Bai and Perron procedure for the German inflation until 1998 and for the euro area inflation after this date conduct to select a maximum of 9 structural breaks for the period and a value of .05 for the trimming parameter ϵ (the number of breaks for the other values of ϵ are 7, 4, 3 and 2 respectively). The average value of inflation which correspond is shown on figures 6 and 7. Two comments could be made. The first is that, when we allow a large number of breaks between 1975 and 2002, we can detect a break in the stochastic dynamics of inflation after the founding of the ECB, even if this break does not coincide with the first months of 1999 (figure 6). In this case, the inflation mean value in the last sample equals 2.41% and 2.36% for 9 and 8 breaks respectively. The second conclusion is that this break is minor in comparison to the past breaks in the German inflation process from 1975 (figure 7). For the 5, 3 and 2-maximum breaks solutions, the mean of the inflation process after 1999 varies between 1.80% and 2.55%. Hence, in conclusion, the Bai and Perron approach concludes on the existence of breaks in the inflation stochastic process.

[20] This solution corresponds to the choice of different newsletters in Kozicki and Tinsley (2001b).

[21] We use the Gauss program of Bai and Perron in this section. The other parameters allow the moment matrix of residuals to be different between subsamples. Results on the number of breaks are similar with the BIC statistic and the modified Schwarz criterion of Liu, Wu and Zideck (1997), except than this latter statistic detect only 6 breaks for the 7-maximum breaks solution.

Figure 1. **German interest rates**

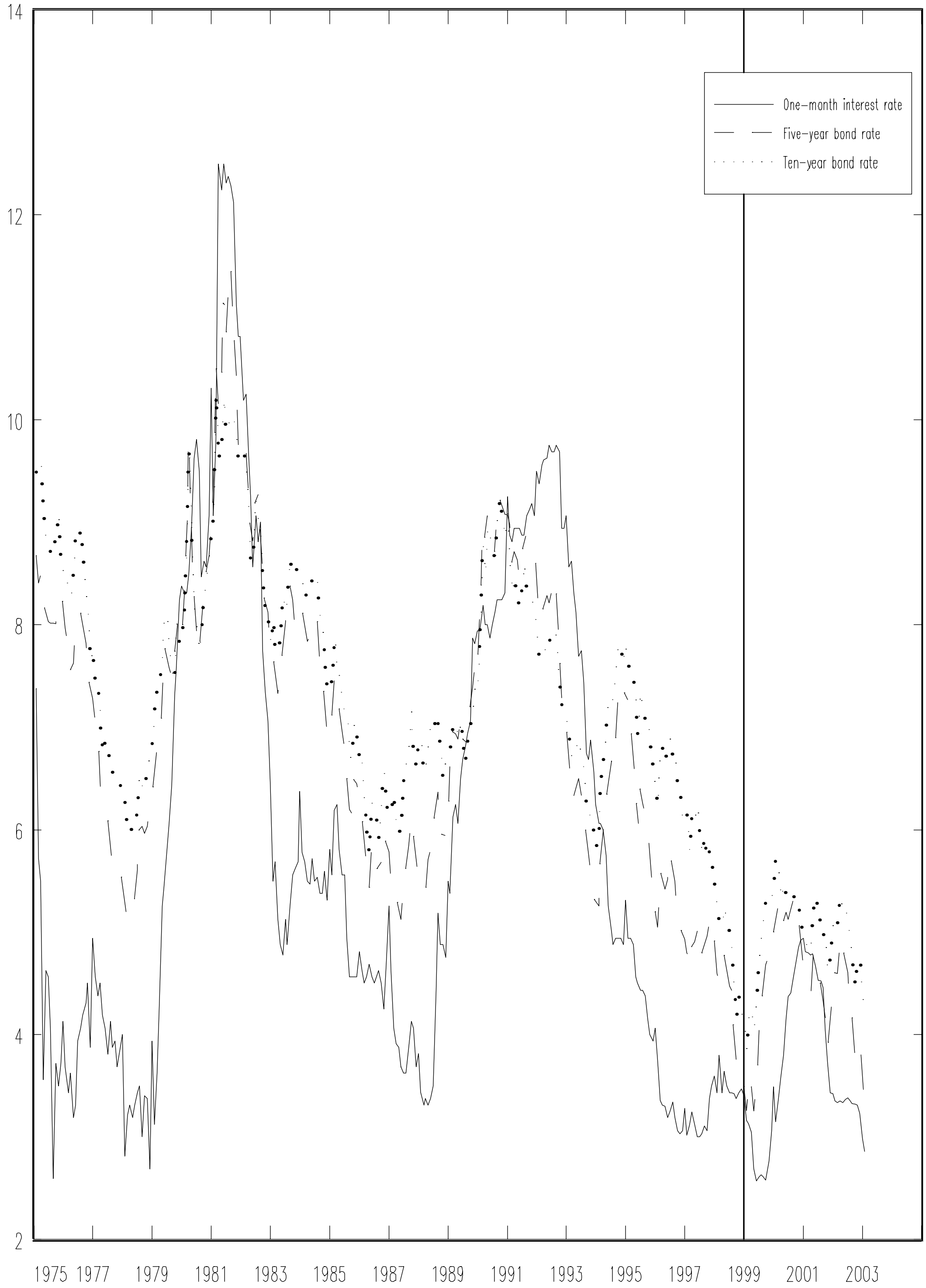


Figure 2. **Volatility of the German long-term interest rates**

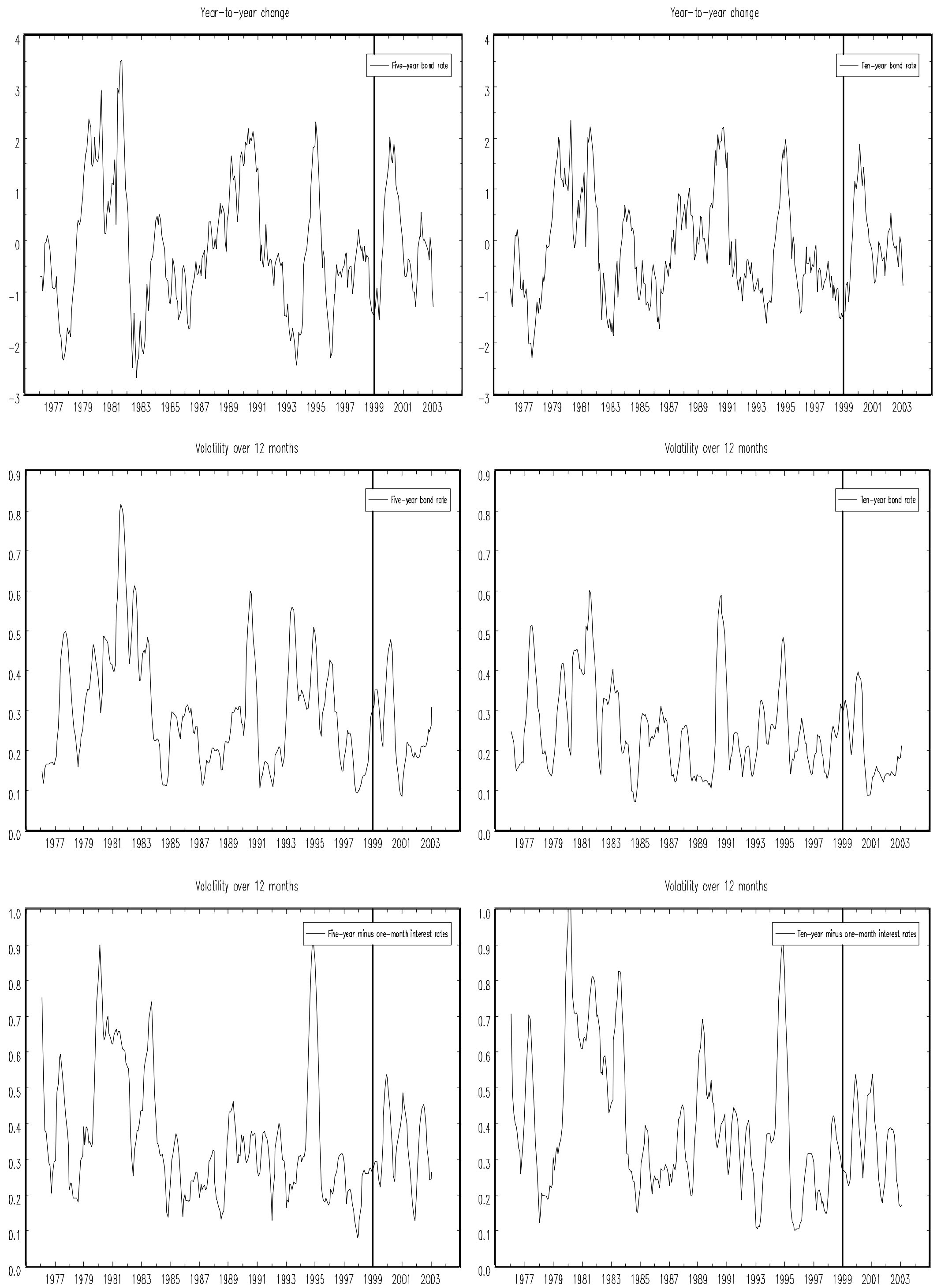


Figure 3. **Nominal short term interest rate endpoints perception**

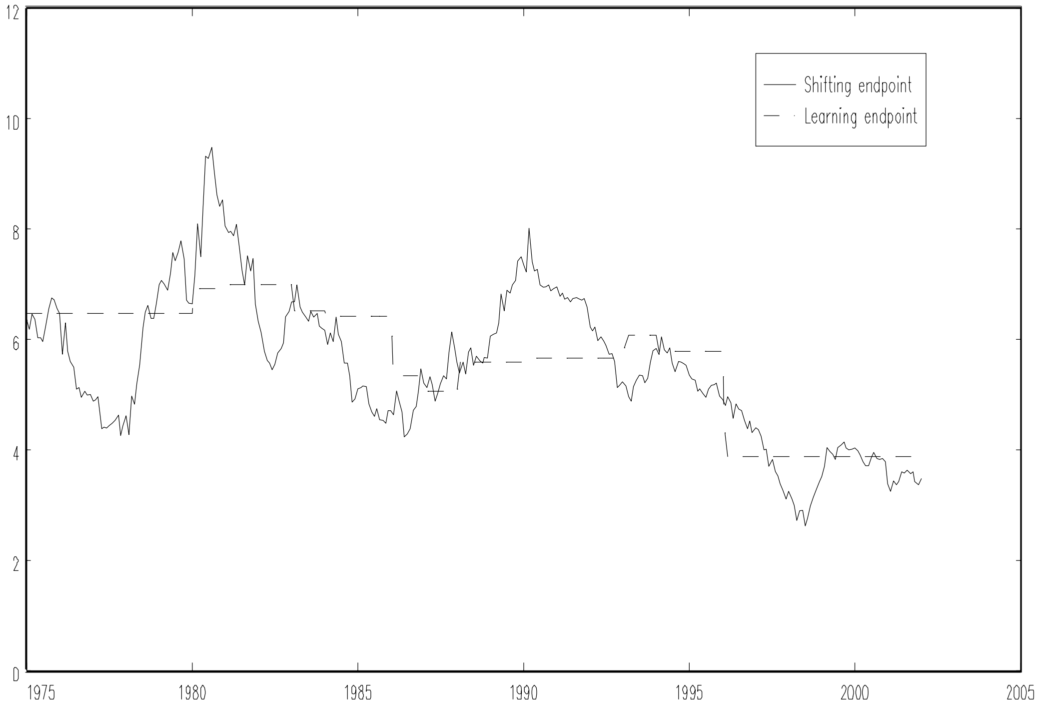
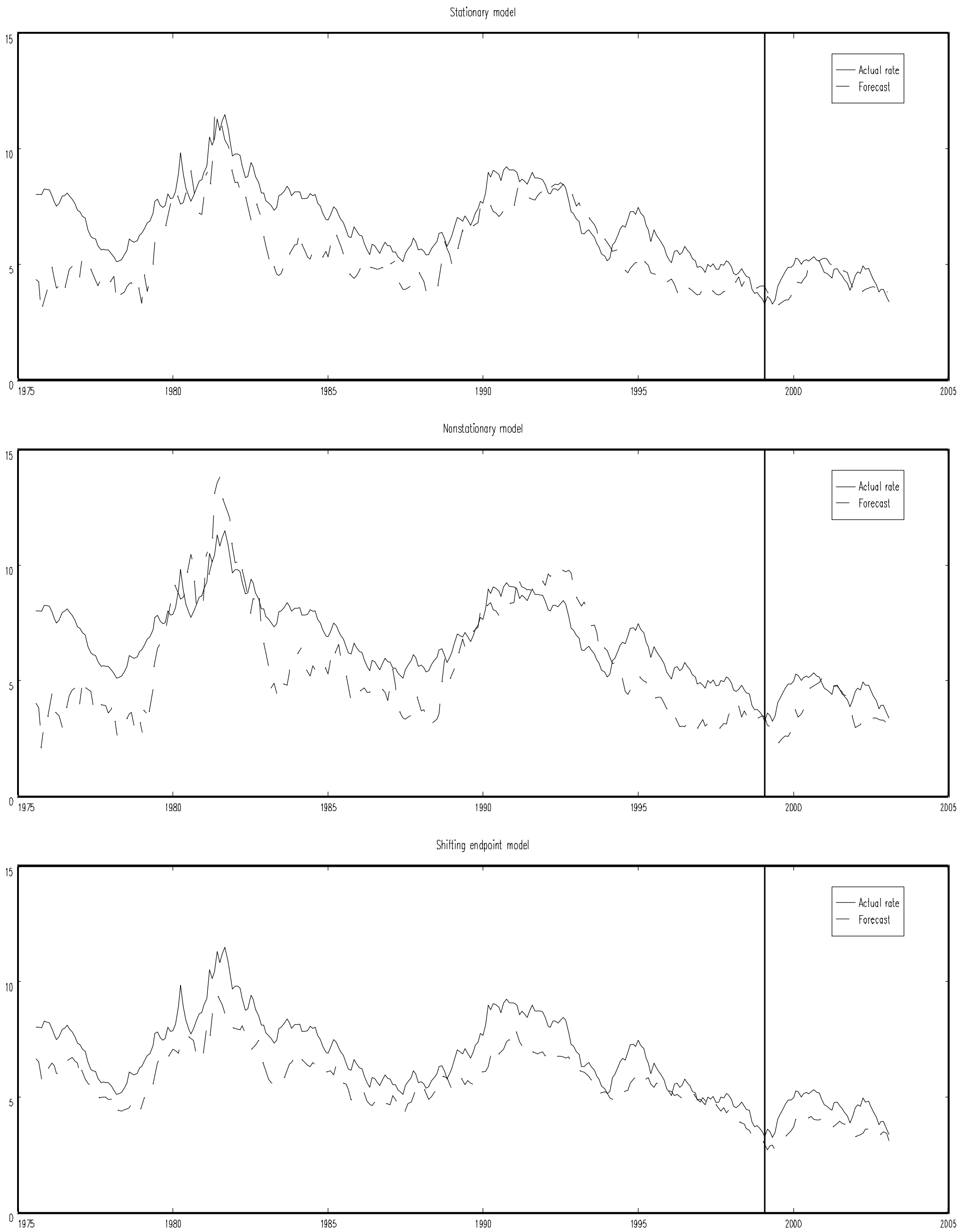


Figure 4. **Expected short rate components of 5-year bond rates**

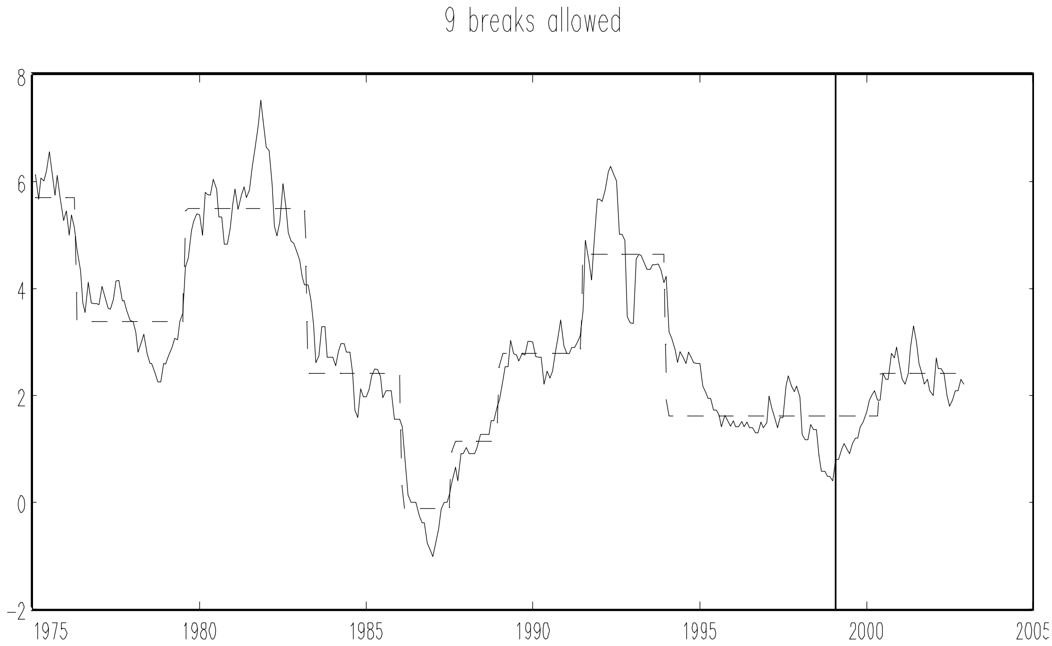


5–year bond rate

Figure 5. **Expected short rate components of 10-year bond rates**

10–year bond rate

Figure 6. **Structural changes in the inflation mean**
Bai and Perron test



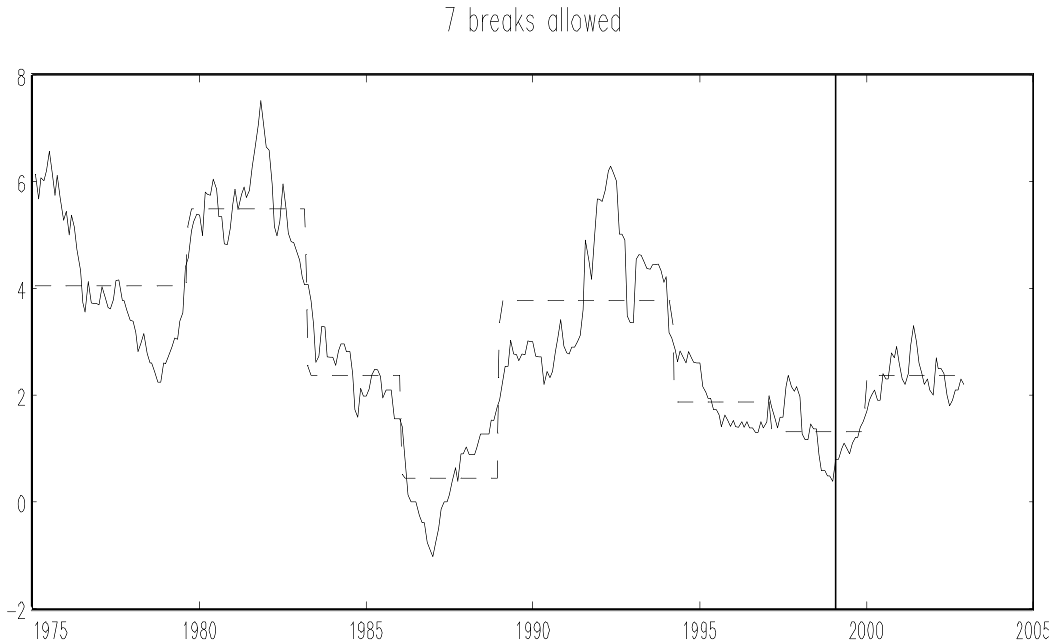


Figure 7. **Structural changes in the inflation mean**
Bai and Perron test

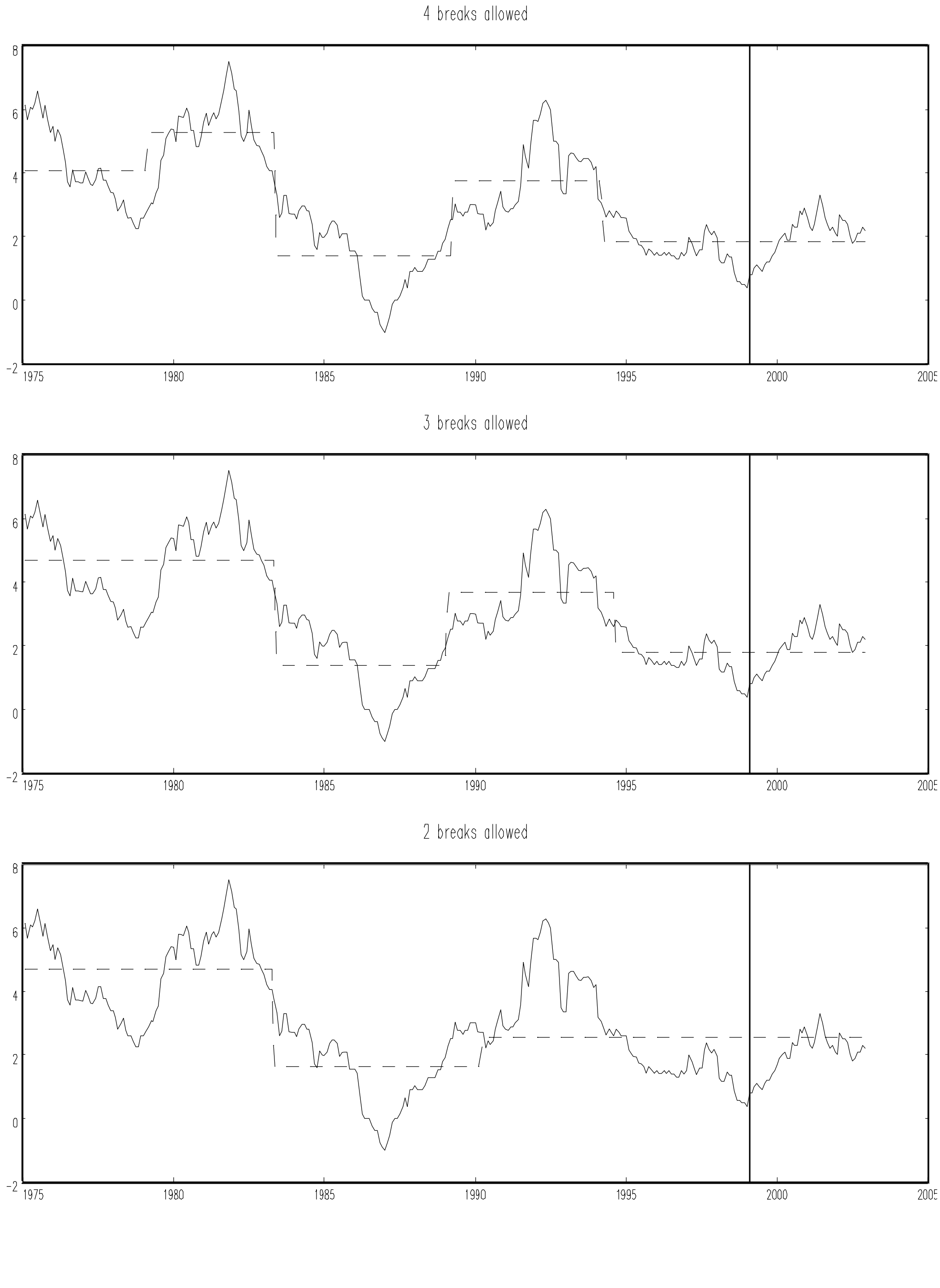


Figure 8. **Learning inflation endpoints obtained from two newsletters**

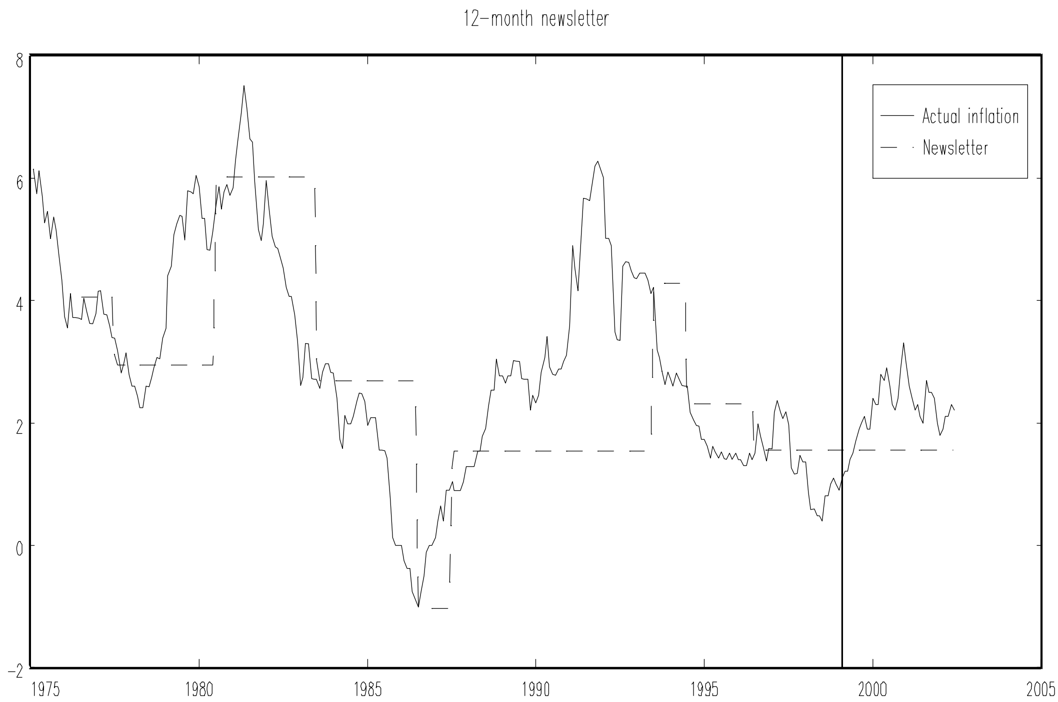


Figure 9. **German and euro area data**

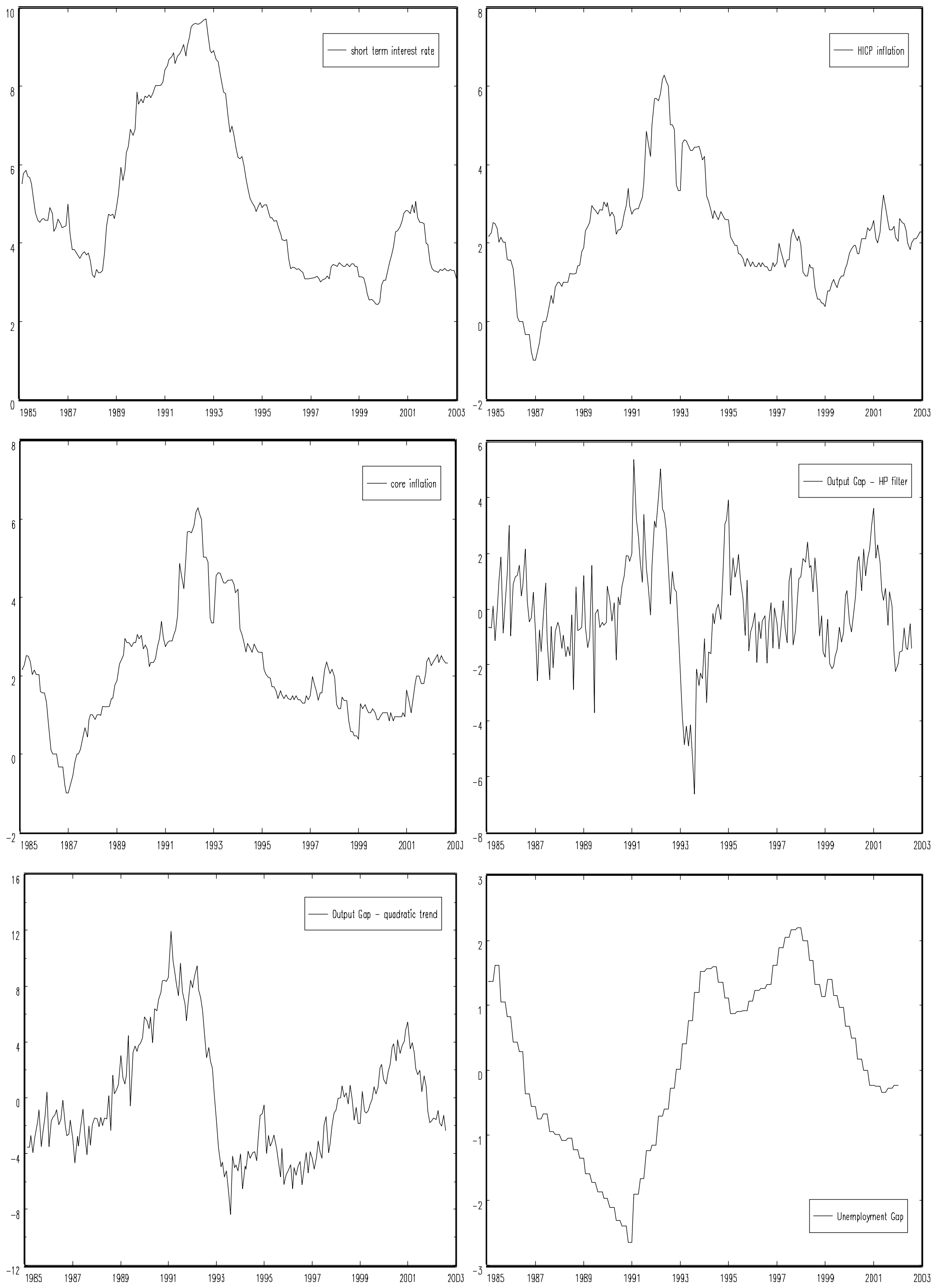


Figure 10. **Nominal short-term interest rate endpoints obtained from reaction functions**

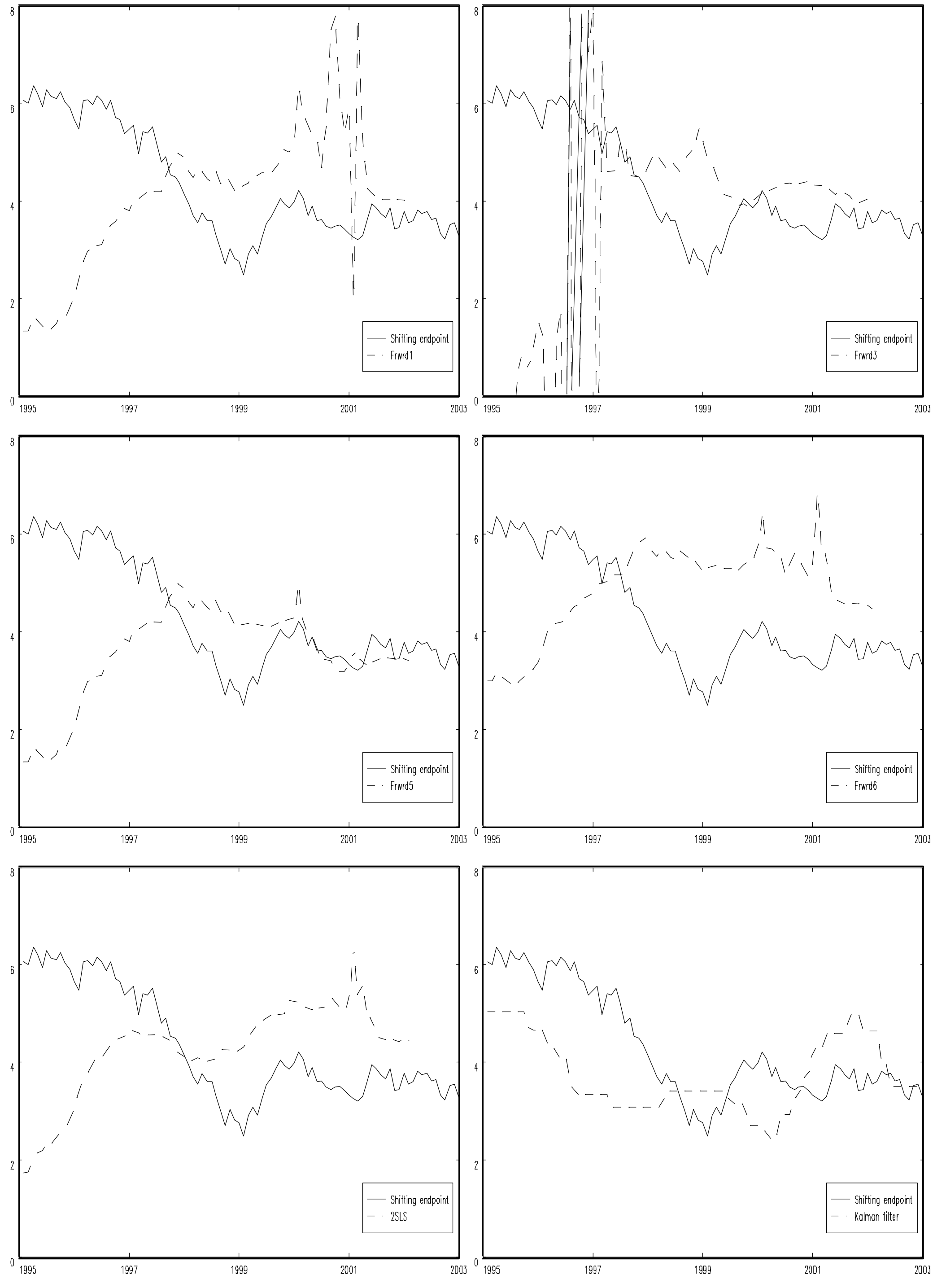
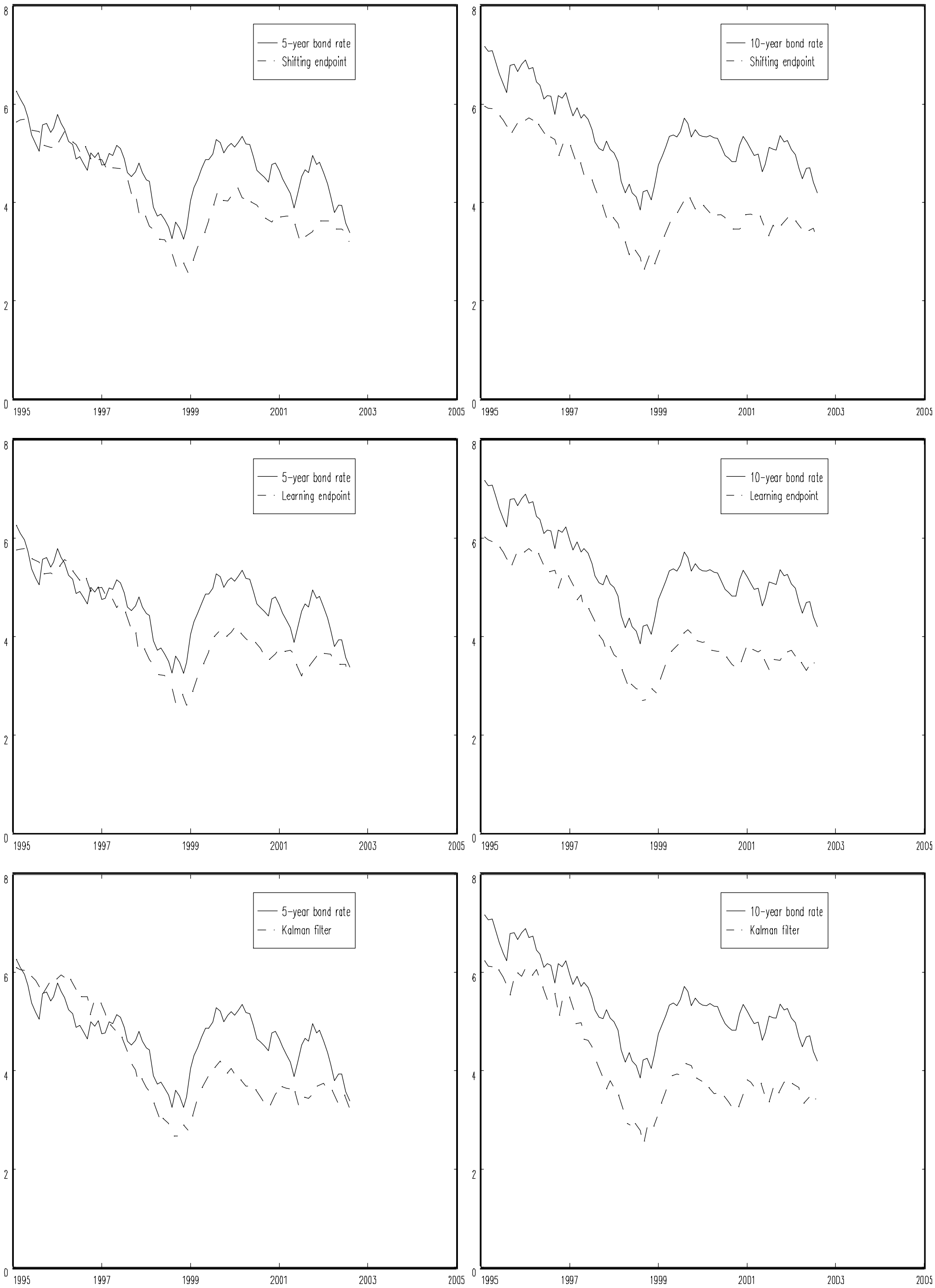


Figure 11. **Expected short rate components of 5 and 10-year bond rates**



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