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The Forward Premium of Euro Interest Rates*

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Abstract

We show that euro forward rates are biased predictors of future interest rates. A small part of this bias arises from unexpected changes in interest rates, while a larger part is explained by the forward premia, which are generally not time-varying. We estimate the the 3-month forward premia for different horizons using forecasts of yields obtained with the Diebold and Li (2006) approach, extended by the inclusion of macroeconomic variables. Confidence intervals for the estimates are computed using a novel bootstrap approach. When using German data for the period before 1999, we detect a break in the dynamic correlation between yield factors, implying that estimates of the euro forward premium using pre-euro data are biased. Although the forward premia of horizons up to 36 months are on average positive, their confidence intervals indicate that they are significantly equal to zero in some periods of time. They are also positively correlated with the ECB policy rate and with a measure of the market perception that future interest rates could be higher than expected.

Key words: expectations hypothesis, interest rates, forecasting, forward premium, euro area.

JEL codes: C32, E43.

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1 Introduction

A popular way of extracting market expectations about future interest rates is the computation of implicit forward rates from the yield curve (Söderlind and Svensson, 1997). However, the forward rate may differ from market expectations by a forward premium. This paper shows that indeed there is a forward premium in the euro area forward rates. There is also a weak evidence that the premium is time-varying. The existence of a premium in the forward rates implies that in order to extract market expectations on interest rates from the yield curve it is first necessary to compute the forward premium.

Estimates of risk premia could be obtained after the estimation of an affine dynamic term-structure model (Duffie and Kan, 1996), which may include macroeconomic variables as additional observable factors (Ang and Piazzesi, 2003). This is the approach employed by Cappiello et al (2006) for the euro area data. The imposition of no-arbitrage ensures that the price of risk is distinguished from the expectations on future yields because it guarantees the existence of a risk-neutral measure. The changes of the risk premia over time incorporated in the affine term-structure models are explained by latent and observed factors, such as macroeconomic variables. As a consequence, the risk premia depend on the factors that explain the yield curve and are not directly available for each forecast horizon that one wants to use the forward rate as forecast. The forward premium for each horizon is computed ex-post based on estimates of the expected short-term interest rate and the forward rate implied by the model (see Appendix of Hordahl et al, 2006).

In this paper, we employ a different approach to compute the forward premia. We employ an extension of the Diebold and Li (2006) approach to obtain forecasts of the yields using all information available at t. The forecasts are consistent with a smooth yield curve and are then employed to compute forward premia for different forecast horizons. We also develop a method for calculating confidence intervals for the forward premia. The evaluation of the significance of the forward premia at each point in time is an important feature of our approach because it helps to decide whether a change in the forward rate is due to a change in the risk premium or instead to a change in expected interest rates. Although the method used in this paper does not impose no-arbitrage, Diebold, Piazzesi and Rudesbush (2005) show that an approach similar to the one of this paper—with two yield factors— is consistent with a two-factor affine no-arbitrage term-structure model under some conditions.

We estimate the forward premia with German zero-coupon interest rates, and also with a shorter sample available after the introduction of the euro, which consists of euro money market rates and swap rates. Although the German data have been exploited before (Hordahl et al, 2006 and Capiello et al, 2006), we are not able to find other papers using only euro data (except when only looking at the money market: Durré, Eujen, Pilegaard, 2003).

When using data before the introduction of the euro with a rolling estimation, we find a large break in the dynamic correlation of yield factors, conditional on macroeconomic variables, in the 2000-2001 period. As a consequence, we suggest that the current forward premia should be estimated with data from 1999 onwards.

Our approach for estimating the forward premia is evaluated by its ability of providing estimates that are compatible with some measures of uncertainty. We show that our estimates of forward premium are positively correlated with the implied bond market volatility in the euro area, but they are more strongly correlated with the key ECB interest rate in the 2000-2001 period. The positive correlation between the forward premium and the official interest rates is only found after the introduction of the euro. This relation might be explained by the fact that agents change their probability distributions about future interest rates when policy rates are increasing (decreasing) such that they increase (decrease) the probability of future interest rates being larger (smaller) than their actual expected values (mean). Because interest rates higher than expected are “bad news”, a move of the distribution to the right implies a higher risk premium. The changes in the skewness of the distribution of the option-implied 3-month EURIBOR futures contracts for the period after 1999 is consistent with this explanation.

The second section of the paper reviews the definitions of risk premium and the expectations hypothesis. The section that follows describes our method to compute the forward premia and their confidence intervals. Section 4 presents empirical results on the biases of the forward rates when forecasting 3-month interest rates. It also shows that this is in part caused by a forward premium. Section 5 presents our estimates of the forward premia, including an evaluation of the instability in the estimation because of changes in the parameters. We also compare our estimates with measures of uncertainty in the financial markets and results from other models in the literature. Section 6 includes some concluding remarks.

2 The Expectations Hypothesis and the Risk Premium

This section presents notation and concepts that are important to understand the different definitions of risk premia. The risk premium on interest rates depends on the definition of the expectations hypothesis of the term structure of interest rates employed to compute the risk neutral component of the interest rates.

Cochrane (2001) defines the *yield* of a bond as "the fictional, constant, known, annual, interest rate that justifies the quoted price of a bond, assuming that the bond does not default" (p. 348). From this definition, the gross yield to maturity of a zero-coupon bond with maturity n and price $P_t^{(n)}$ is $Y_t^{(n)}$ that satisfies $P_t^{(n)} = \frac{1}{\left[Y_t^{(n)}\right]^n}$. Assuming that the price of a bond at the maturity date is 1 ($P_t^{(0)} = 1$), the return of holding a n-period bond until maturity (return to maturity) corresponds to $R_t^{(n)} = \frac{1}{P_t^{(n)}}$. Using these variables in logs, we can see that the log-yield of a zero-coupon bond ($y_t^{(n)} = -\frac{p_t^{(n)}}{n}$) corresponds to the log-return per period ($y_t^{(n)} = \frac{r_t^{(n)}}{n}$).[1] In addition to the yield and the return to maturity, the holding period return

[1] In the Cochrane (2001) notation, the yield (and the return) corresponds to one plus the interest rate. In

and the forward rate are also obtained from the log-price of a zero-coupon bond. The *holding period return* is the return from holding a bond with maturity n over the next period (i.e, from t to $t+1$): $hpr_{t+1}^{(n)} = p_{t+1}^{(n-1)} - p_t^{(n)}$. The *one-period forward rate* is the interest rate contracted today to start in n periods from now and to be paid back in $n+1$ periods from now, that is, $f_t^{(n,n+1)} = p_t^{(n)} - p_t^{(n+1)}$. The one-period forward rate can be also written using the yields as $f_t^{(n,n+1)} = y_t^{(n+1)} + n(y_t^{(n+1)} - y_t^{(n)})$.

"The yield curve is a plot of yields of zero-coupon bonds as a function of their maturity" (Cochrane, 2001, p. 352). Most of the times, the yields of zero-coupon bonds are increasing with maturity. Based on the definition of the forward rate using the yields, one can show that the one-period-forward rate is above the yield for the same n if the yield curve is upward slopping.

The expectations hypothesis describes the relationship between returns on bonds of different maturities. It is based on the idea that expectations about future interest rates affect the current level of long rates. As Cox, Ingersoll and Ross (1981) point out there are four different formulations of the expectations hypothesis:

(*i*) The yield of a zero-coupon bond that matures in n periods is equal to the average of the expected one-period yields (yield to maturity hypothesis - YTM):

$$y_t^{(n)} = \frac{1}{n} E_t(y_t^{(1)} + y_{t+1}^{(1)} + ... + y_{t+n-1}^{(1)})$$

(ii) The return to maturity of a n-period bond is equal to the expected return of rolling over a series of single-period bonds (return to maturity hypothesis - RTM):

$$r_t^{(n)} = E_t(r_t^{(1)} + r_{t+1}^{(1)} + ... + r_{t+n-1}^{(1)})$$

(iii) The one-period forward rate n-periods ahead is equal to the corresponding expected one-period spot rate (unbiased expectations hypothesis - UEH):

$$f_t^{(n,n+1)} = E_t(y_{t+n}^{(1)})$$

(iv) The expected holding period return of a bond with maturity n is equal to the current one-period interest rate (local expectations hypothesis - LEH):

$$E_t(hpr_{t+1}^{(n)}) = y_t^{(1)}$$

In practice the above relations do not necessarily hold. The difference between the left-hand side and the right-hand side of the above equations is then called the *risk premium.* If the expectations hypothesis holds in its pure form, the risk premium will be zero. However, it

logs, this distinction is not relevant since $\ln(1+i) \approx i$. It is also important to notice that $y_t^{(n)}$ corresponds to the continuously compounded interest rate. This is so because the relation between the continuously compounded interest rate (i^c) and the price of a bond with maturity n is $e^{ni^c} = \frac{1}{P_t^{(n)}}$, and the relation between i^c and an annually compounded interest rate (i^a) is $i^c = \ln(1+i^a)$.

is generally considered that the expectations hypothesis holds if the risk premium is constant over time. For the empirical testing of the expectations hypothesis, the definition (i) is the most popular, because it can be tested using restrictions in a vector autoregression of yields of different maturities and/or using cointegration. The results depend in general on how large n is. For large n, the hypothesis of a constant risk premium is most of times not valid. Another way of testing the expectation hypothesis is to use a predictive regression to verify the definition (iii). Results of this type of tests are presented in section 4.4. Definition (iv) is used when imposing "no-arbitrage conditions" in the context of building dynamic term structure models (Singleton, 2006).

The four different formulations of the expectations hypothesis imply different measurements of the (time-varying) risk premium. The premium derived from definition (i) is frequently called yield premium, term premium or rollover term premium. The premia arising from definitions (iii) and (iv) are called (one-period) forward premium and (one-period) holding premium. Finally, the premium corresponding to definition (ii) is not usually calculated. The nomenclature of the different risk premia is sometimes confusing. In fact, many authors (as, for example, Singleton, 2006) use the name term premium to refer generically to risk premium because it arises from the existence of different maturities in the yield curve. In this paper we will refer to the premia derived from definitions (i) up to (iv) respectively as yield premium, return premium, one-period forward premium and one-period holding premium.

In Appendix 1, we show that these four different ways of defining the expectations hypothesis are mathematically equivalent. As a consequence the yield premium of a bond with maturity n is equal to its return premium divided by n, to the average of the one-period forward premia, and to the one-period holding premia. However these are particular results for the case when the variables are in logs and time is discrete. Cox, Ingersoll and Ross (1981) show that, in continuous time and when the previous versions of the expectations hypothesis are levels, only the UEH and YTM are equivalent.[2] When the four forms of the expectations hypothesis are analysed in discrete time and in levels, they are in general different from each other, except that the UEH is equivalent to the RTM if the levels of future interest rates are uncorrelated. As a consequence, bond prices, yields and the risk premia generated by the different formulations will be mutually inconsistent. Anderson et al (1996) show that, when time is continuous and the variables are in levels, the instantaneous forward premia, which is zero by definition in the case of UEH, is negative if the LEH is valid and positive if RTM is valid.

[2] This results from the Jensen´s inequality which states that for a random variable X, if $g(X)$ is a strictly convex function, then $E\left[g(X)\right] > g\left[E(X)\right]$ and that the opposite is true when $g(X)$ is a strictly concave function.

3 Procedures for the Estimation of the Forward Premia and their Confidence Intervals

In the previous section, we described four different ways of writing the expectations hypothesis and their implications for the definition of the risk premia. In this section, we describe how an extension of Nelson and Siegel (1987) approach to fit a yield curve can be applied for the computation of forward premia for all maturities and forecast horizons. The method will be applied specifically for the forward premium, but it could also be useful to compute the other definitions of risk premia. The emphasis on the forward premium arises from the fact that we want to use the information in the yield curve to obtain market forecasts of future interest rates. We also propose a bootstrap procedure to compute confidence intervals for the forward premium. The confidence intervals are useful to evaluate whether the forward premium at a specific point in time is statistically different from zero and also to verify the uncertainty around the computation of the forward premium.

The method presented here for the computation of the premia is an extension of Diebold and Li (2006), which is based on the Nelson and Siegel (1987) parametric yield curve fitting. The first step of the procedure applies the Nelson and Siegel (1987) regression for computing yield factors. These factors are then modelled jointly with macroeconomic variables in a vector autoregression (VAR). The estimates of the VAR are used to compute forecasts of the yields factors up to $t+h$ using information at t. Using factor forecasts, yield forecasts are obtained with the Nelson and Siegel regression. At date t, the estimate of the τ-period forward premium for horizon h is the difference between the implicit forward rate of maturity τ contracted in t to start in $t+h$ and the forecasts at t for the yield of maturity τ in $t+h$.

Diebold and Li (2006) show, with US data, that their approach generates forecasts of interest rates that are more accurate than standard benchmark forecasts, including the ones obtained with the Cochrane and Piazzesi (2005) forward curve regression. In addition, in comparison to accurate forecasts of simple statistical models, the advantage of the Diebold and Li (2006) approach is that the forecasts are consistent for rates of each maturity, so that the implied yield curve is smooth and the forward rates are positive. Favero and Kaminska (2006) in their comparison of estimates of yield premium also use the Diebold and Li (2006) forecasts for computing the premium. The approach of this paper in comparison to Favero and Kaminska (2006) presents the following differences: (a) a three-factor yield curve model is employed instead of a two-factor model; (b) macroeconomic variables are included in the VAR; and (c) confidence intervals for the estimates of the forward premia are computed.

3.1 The τ-period Forward Premium

In general one might be interested in forecasting interest rates of maturity higher than one period. At date t, the $\tau-$period forward rate, i.e., the interest rate contracted in t to start in period $t+n$ and with maturity in period $t+n+\tau$, corresponds to:

$$f_t^{(n,n+\tau)} = \frac{1}{\tau}\left[\tau y_t^{(n+\tau)} + n(y_t^{(n+\tau)} - y_t^{(n)})\right] = \frac{1}{\tau}\left[(n+\tau)y_t^{(n+\tau)} - ny_t^{(n)})\right]. \tag{1}$$

Thus the $\tau-$period forward premium is:

$$frp_t^{(n,n+\tau)} = f_t^{(n,n+\tau)} - E_t(y_{t+n}^{(\tau)}). \tag{2}$$

The forward rates $f_t^{(n,n+\tau)}$ can be computed using the yield (spot) rates $y_t^{(n+\tau)}$ and $y_t^{(n)}$. However, the maturities of observable yields may not match the ones required to compute the forward rates that we are interested in. Thus, it is necessary to fit a curve for the observable yields, so that one can use the fitted yields to compute forwards for any desired maturity and horizon. For the estimation of the $\tau-$period forward premium, one also needs $E_t(y_{t+n}^{(\tau)})$, which can be estimated by a n-period ahead forecast for the yield with maturity τ ($\hat{y}_{t+h|t}^{(\tau)}$, where $h = n$).

3.2 Fitting and Forecasting the Yield Curve

Nelson and Siegel (1987) propose a parametric approach for fitting the yield curve. The yield curve is fitted for a group of observed zero-coupon bonds with maturities τ at a given point in time $y_t^{(\tau)}$. The Nelson and Siegel equation for the spot (yield) rate is:

$$y_t^{(\tau)} = \beta_{1t} + \beta_{2t}\left(\frac{1-e^{-\theta_{1t}\tau}}{\theta_{1t}\tau}\right) + \beta_{3t}\left(\frac{1-e^{-\theta_{1t}\tau}}{\theta_{1t}\tau} - e^{-\theta_{1t}\tau}\right), \tag{3}$$

where β_{1t}, $\beta_{1t}+\ \beta_{2t}$ and θ_{1t} must all be positive. The parameters β_{1t}, β_{2t} and β_{3t} are called yield factors and are interpreted as the level (L_t), the symmetric of the slope ($-S_t$) and the curvature (C_t) of the yield curve. θ_{1t} is the parameter that measures the rate of the exponential decay of the loading of the second and the third factors. Smaller θ_{1t} implies slower decay. This parameter also defines the maturity at which β_{3t} has larger weight. Diebold and Li (2006) fix θ_{1t} in the value that the maturity of almost 3 years has the highest loading for β_{3t}. An advantage of keeping θ_{1t} fixed is that the factors can be estimated by the usual least squares formula. Diebold and Li (2006) also argue that the estimates of the yield factors are more stable over time when θ_{1t} is fixed, which is an advantage when one is interested in predicting them.

Svensson (1994) suggests the inclusion of an additional term to increase the flexibility and to improve the fit of the yield curve:

$$y_t^{(\tau)} = \beta_{1t} + \beta_{2t}\left(\frac{1-e^{-\theta_{1t}\tau}}{\theta_{1t}\tau}\right) + \beta_{3t}\left(\frac{1-e^{-\theta_{1t}\tau}}{\theta_{1t}\tau} - e^{-\theta_{1t}\tau}\right) + \beta_{4t}\left(\frac{1-e^{-\theta_{2t}\tau}}{\theta_{2t}\tau} - e^{-\theta_{2t}\tau}\right). \tag{4}$$

β_{4t} and θ_{2t} have the same characteristics of β_{3t} and θ_{1t}, respectively, and θ_{2t} is also restricted to be positive. These additional parameters help capturing an additional hump- or u-shape in the yield curve. Because parsimonious models are more successful for out-of-sample forecasting, it is not clear that the additional term improves yield forecasts. Both approaches by Svensson

(1994) and the Nelson and Siegel (1987) are popular in the modelling the yield curve by central banks (see BIS, 2005). In the empirical exercise, we will evaluate the fit of the yield curve with both approaches, and also the loss of fit of keeping θ_{1t} fixed.

The Nelson and Siegel (1987) approach for fitting the yield curve can be used to forecast yields of different maturities (Diebold and Li, 2006). Using the name of factors and a fixed θ_{1t}, the Nelson and Siegel regression for forecasting a yield of maturity τ at h-steps ahead conditional on information at t is:

$$\hat{y}^{(\tau)}_{t+h|t} = \hat{L}_{t+h|h} - \hat{S}_{t+h|t}\left(\frac{1-e^{-\bar{\theta}_1\tau}}{\bar{\theta}_1\tau}\right) + \hat{C}_{t+h|t}\left(\frac{1-e^{-\bar{\theta}_1\tau}}{\bar{\theta}_1\tau} - e^{-\bar{\theta}_1\tau}\right). \tag{5}$$

Diebold and Li (2006) suggest the estimation of an AR(1) for each factor to be able to compute $\hat{L}_{t+h|t}, \hat{S}_{t+h|t}$ and $\hat{C}_{t+h|t}$. However, there is some important dynamic correlation between the slope, the level and the curvature. Thus, we consider a VAR(1) to be more adequate. Using the estimated factors at each time $t = 1, ..., T$, we define a VAR(1) for modelling the vector $x_t = \begin{pmatrix} \hat{L}_t & -\hat{S}_t & \hat{C}_t \end{pmatrix}'$ as:

$$x_t = c + \Phi_1 x_{t-1} + \varepsilon_t. \tag{6}$$

Conditional on the estimated parameters, we use this VAR to generate h-step-ahead forecasts as:

$$\hat{x}_{t+h|t} = (1 + \hat{\Phi}_1 + ... + (\hat{\Phi}_1)^{h-1})\hat{c} + (\hat{\Phi}_1)^h x_t. \tag{7}$$

There are several recent papers modelling the relation between factors of the yield curve and some important macroeconomic variables (for example, Ang and Piazzesi, 2003; Diebold et al, 2006; Rudebush and Wu, 2004; Hordahl et al, 2006). One simple way of adding information of macroeconomic variables to predict the factors of the yield curve is to augment the VAR of equation (6) with a small group of variables. We estimate two specifications of the VAR: one with only the yield factors and another including also inflation and real activity growth. These macroeconomic variables were chosen because there is evidence in the literature of a strong dynamic relation between them and yield factors. For the euro area or some euro area countries, predictive regressions show that the slope helps predicting real activity growth and inflation (see, for example, Estrella et al (2003), Moneta (2003) and Duarte et al (2005)). Hordähl et al (2006) estimate an affine term-structure model extended with structural equations for inflation, output gap and the short-term interest rate with Germany data, and they conclude that macroeconomic variables help predicting the yield curve, even though the yield curve does not provide useful additional information in forecasting macroeconomic variables. In particular, their results suggest that inflation and output gap affect the curvature of the yield curve. Cappiello et al (2006) estimate the model of Hordähl et al (2006) for the pre- and post-euro periods and they conclude that the relevance of macro factors in explaining the behaviour of the risk premium does not change with the introduction of the euro.

An alternative to our approach of estimating the factors in a first step and the dynamic relation of the factors in the second step is the one proposed by Diebold et al (2006). Diebold et al (2006) show how to jointly estimate the yield factors (eq. 3) and the coefficients of a VAR of yield factors and macroeconomic variables (eq. 6), using a state-space representation by Kalman filtering and maximum likelihood estimation. A disadvantage of the joint estimation of the parameters and the three unobserved factors, which are non-linearly related with observable yields, is the challenge in the numerical optimization procedure. In addition, because the yields have a high persistence over time, the time-dependence of the factors may be captured even when the regression is computed independently each time. Another issue is that the three factors explain most of the variation of the yields, implying that the inclusion of the macro variables does not affect the estimation of the factors. However, when using the VAR for forecasting yield factors, it is important to consider the dynamic relation between the macro variables and the yield factors because there is a strong relation between them. With the support of previous statements, our two-step estimation may generate yield forecasts similar to the ones implied by the model of Diebold et al. (2006). In addition, the use of method which is less demanding in computation reduces the problem of using a sample as short as the one available after the euro introduction. Similar two-step approaches have been employed by Carriero et al (2006) and Favero and Kaminska (2006).

3.3 Confidence Intervals for the Forward Premium

Based on forecasts of the yields $\hat{y}^{(\tau)}_{t+h|t}$ and on the implied forward rates, the forward premium can be computed with equation (2). However, the computation of the forward premia does not give information on whether they are statistically different from zero at each point in time. Even if the rejection of the expectations hypothesis indicates the existence of a time-varying forward premia, it may be the case that at a specific point in time the forward premia may not be statistically significant from zero. We design a bootstrap procedure to compute confidence intervals for the forward premia at each specific date for which forecasts for the short-rate are obtained from the forward rates. If the confidence intervals do not include zero, one should not use the forward rates for forecasting short-term interest rates without taking the forward premium into account.

The forward premium is computed with an estimate of the conditional mean of future yields $\hat{y}^{(\tau)}_{t+h|t}$. The estimates of the residuals of the VAR of the yield (and macro) factors provide information on the type of unexpected shocks that may affect the shape of the future yield curve. These shocks are the main source of uncertainty of the yield forecasts. The information on past shocks can be used to compute the conditional density of future yields, so we can assess the uncertainty in the computation of $\hat{y}^{(\tau)}_{t+h|t}$ and, as consequence, of the forward premium.

Without making any assumptions on the shape of the predictive density, one can compute the density using bootstraps from the residuals of the estimated VAR: $\hat{\varepsilon}_t = x_t - \hat{c} - \hat{\Phi}_1 x_{t-1}$. At each replication, one can bootstrap a sequence of shocks of size h: $\hat{\varepsilon}^*_1, ..., \hat{\varepsilon}^*_h$. Conditional on information at t and with the estimates of the VAR coefficients, one can compute $\hat{x}^*_{t+1|t}$

using the first shock in the first bootstrap sequence, then recursively and adding one shock at a time, one can get $\hat{x}^*_{t+2|t}, ..., \hat{x}^*_{t+h|t}$. Using the information on future yield factors containing in $\hat{x}^*_{t+1|t}, ..., \hat{x}^*_{t+h|t}$, a sequence of predicted $\tau-$yield $\hat{y}^{(\tau)*}_{t+1|t}, ..., \hat{y}^{(\tau)*}_{t+h|t}$ is computed using the Nelson and Siegel equation. Then, if we compute m sequences of h-step ahead forecasts of the $\tau-$yield, one has an empirical predictive density for each h-step forecast and for the forward premium (recall $\widehat{frp}^{*,(n,n+\tau)}_t = \hat{f}^{(n,n+\tau)}_t - \hat{y}^{*,(\tau)}_{t+h|t}$). The 5 per cent and 95 per cent empirical quantiles are then used to compute the 90 per cent confidence interval for the forward premium. The design of the bootstrap guarantees that the intervals are centered.

Note that in this procedure, we are just accounting for the unexpected shocks that may affect future values of the yield factors. Carriero et al (2006) take into account the uncertainty on the estimation of the VAR (that is, on the coefficients of eq. 6) to compute intervals for the long-term interest rate implied by the expectations hypothesis. Our assumption is that in comparison with the measurement error of using the Nelson and Siegel (1987) approach to fit the yield curve and the uncertainty on the estimation of the dynamics between the yields (and macro) factors, the unexpected shocks on the yield curve are more important to measure the uncertainty on the computation of the forward premium. In this context, more information inside the VAR, such as additional macro variables reduces the imprecision in the estimation of the forward premium.

4 Is there a Forward Premium?

In this section, we find that there is a significant positive forward premium in the euro 3-month forward rates. The evidence that the forward premium is time-varying is weak.

4.1 Data Description

We use three different sets of data of continuously compound zero-coupon interest rates sampled monthly. These sets of data are described in more detail in Table 1.

The first one ("data 1") is the longest (1986:1-2006:6) and consists of German data: Libor interest rates and zero-coupon yields estimated by the Bundesbank. We choose the start of the sample in 1986 to exclude the effect of changes in the money market operating procedures of the Bundesbank in February 1985. There are several arguments in favour of using German data. As pointed out in Cappiello et al (2006), the German bond market has to some extent been a benchmark for European bond market as a whole. In addition, Germany did not observe currency crisis in the ERM period, as a consequence, the data are less affected by intra-area currency effects. The possibility of using a "synthetic euro" term structure is not considered because the differences in macroeconomic environment and monetary policy in the different euro area countries, imply that the aggregation of the data is senseless. This argument is to some extent also valid for the period after 1999 because it may not be meaningful to model the yield curve using data on "synthetic bonds" that are not traded. For the later period an alternative is to use data on the euro swap curve. To be consistent, we prefer to use only German bond yield data in "data1".

The other two sets of data consist of euro interest rates, so they have only information after 1999. "Data 2" use EURIBOR rates for maturities up to 12 months and swap rates for the longer maturities. "Data 3" contain estimates of zero-coupon rates for all maturities, which are based on money market and swap rates and include maturities for each 3 months up to 10 years. "Data 1" and "data 2" have maturities only at each 3 months up to 1 year and at each year up to 10 years.

The macroeconomic series are for Germany in the pre-euro period and for the euro area since 1999. In addition to our interest in the behaviour of euro area, another reason for using euro area data even when the interest rates are for Germany ("data 1") is that monetary policy, which is an important determinant of the yield curve, is based on the euro area macro behaviour after 1999.

4.2 Fit of the Estimated Yield Curve and Evaluation of Forward Rates as Forecasts

Two parametric methods for fitting the yield curve using three (Nelson and Siegel, 1987) and four factors (Svensson, 1994) are discussed in section 3. Some authors (Favero and Kaminska, 2006) have also argued the use of two factors. Our preliminary analysis indicate that two-factor regressions are unable of characterising intermediate maturities (say 1 and 2 years). Table 2 presents the square root of the mean of the squared residuals for each maturity of estimating a yield curve with three and four factors for each month in our data (equations 3 and 4).[3] Table 2 also presents the fit when θ_{1t} is kept fixed and equal to 0.0542, implying that the curvature has the largest loading at a three-year maturity. One can see why Svensson's approach is so popular: the average of the residuals for a specific maturity over time is not larger than 4 basis-points. The larger negative impact of the exclusion of one factor is in general on the adjustment of the shortest maturity. When θ_{1t} is fixed, the largest losses of fit are detected for the one-month maturity and some intermediate maturities (9 months and 1 year). Even in worse fit scenario (NSf), the maximum error is only of 8 basis-points. The best fit is obtained with the sample that includes more maturities ("data 3"). This is clearer in the case of the Svensson specification, which has more parameters that need estimation.

If one is not sure whether the loss of fit of using Nelson and Siegel (1987) with fixed loading instead of Svensson (1994) is important, the results in Table 3 indicate that the choice between 3 and 4 factors and between fixed and time-varying loadings is not relevant for using forward rates to predict the 3-month interest rates. Based on the estimated yields and using expression (1), the 3-month forward rates ($\tau = 3$) are computed quarterly for horizons up to three years (n corresponds to $h = 3, 6, ..., 36$). Both the forward rates and the observed 3-month interest rates $y_{t+h}^{(3)}$ are transformed to be annually compounded in order to be compatible with the usual form

[3] The estimates of the curve with both approaches are obtained with Gauss CML package (with BFGS) imposing restrictions on θ_{1t} and θ_{2t}, so that the economic meaning of the estimated factors is kept.

of presentation of these series. The emboldened average forecast errors of Table 3 indicate that the t-statistic calculated using the estimates and standard errors is larger than 2, implying a rejection of the null that the forecast error is equal to zero. The rejection of the null in favour of a positive forecast bias does not generally depend on the method to fit the yield curve, but it depends on the sample. The shorter sets of data indicate more frequent rejections of the null and larger positive bias than "data 1".

4.3 Forecasts with the Extended Diebold and Li (2006) Approach

The bias identified when forward rates are directly used as forecasts of future spot rates may be the result of either a forward premium or of poor forecasts of the future short-term interest rates, due to unexpected shocks. As a consequence, we evaluate, in this section, whether the Diebold and Li (2006) approach augmented by macroeconomic variables can generate forecasts of short-term interest that are unbiased. If there is still a failure in predicting at long horizons, we would attribute it to errors in forecasting interest rates. In fact, unexpected changes in interest rates take longer to be incorporated when forecasting long horizons because of the longer time required to observing them.

For computing the forecasts, we fix $\bar{\theta}_1$ (eq. 5), as in the previous section. The VAR (eq. 6), which is used to generate h-step-ahead forecasts of the 3-month interest rate, is computed using only the information of yield factors ($x_t = (\ \hat{L}_t \quad -\hat{S}_t \quad \hat{C}_t\ )'$) and also adding information of real activity growth (g_t) and inflation (π_t) ($x_t = (\ \hat{L}_t \quad -\hat{S}_t \quad \hat{C}_t \quad g_t \quad \pi_t\ )'$).

Table 4 presents the t-statistics of the tests of Granger-causality to check the dynamic relationship between yield factors and macroeconomic variables. In the case of inflation, we can only find evidence that it helps forecasting the yield curve with "data 2". There is some evidence that growth helps predicting the yield curve and that the yield curve—the curvature and to a lesser extend the slope— helps predicting growth. Even though the addition of macroeconomic variables will probably improve forecasts of the yield curve, this improvement may be small because of the contribution of those variables not being widespread across sample period and factor.

Table 5 presents the mean of the forecast errors in predicting the 3-month interest rates at horizons between 3 and 36 months. The forecasts use full sample information on the estimation of the VAR parameters, but they use information on yields up to t. The emboldened entries indicate again the rejection of the null that the average forecast error is equal to zero. There is some evidence of bias at long horizons and short samples ("data 2" and "data 3"), but the bias is on average three to five times smaller than using the forward. There is also some weak evidence that the inclusion of macroeconomic variables improves the forecasting performance.

Because Table 5 uses estimates of sets of data with different sample sizes, we check whether there are differences across them when using the same sample size. Table 6 presents the mean of the forecast errors for all sets of data, but only using data from 2000:2 onwards to compute the VAR parameters. The size of the biases are similar across sets of data and they are significant different from zero in long horizons. The inclusion of macroeconomic variables reduces marginally the biases.

Summarizing the results for the period after 1999, they indicate that our approach generates unbiased forecasts at short and medium horizons and that the forward rates generate biased forecasts for horizons larger than 3 months. This suggests that the bias obtained when using forward rates as forecast is in part caused by a forward premium. For long horizons, part of the bias arises from unexpected changes in interest rates. The estimates of the biases at the 3-year horizon presented in Table 6 suggest that 1/4 of the bias incurred by the use of forward rates as forecasts is caused by unexpected shocks, while 3/4 of the bias could be explained by forward premium.

4.4 Test of the Expectations Hypothesis

The previous results on the bias of forecasts when using forward rates and the extension of the Diebold and Li (2006) approach indicate that there is a forward premium in the period after the introduction of the euro. We now apply formal tests of the expectations hypothesis using regressions that allow us to check explicitly whether the forward premia are time-varying .

The standard test of expectations hypothesis with forward rates uses the following predictive regression:

$$y_{t+h}^{(3)} - y_t^{(3)} = \beta_0 + \beta_1 (f_t^{(h,h+3)} - y_t^{(3)}) + \varepsilon_{t+h}.$$

The inclusion of the current short-term interest rate in both sides of the regression reduces bias in the estimates that may arise from non-stationarity of the variables. If the expectations hypothesis is valid and there is rational expectation, the restrictions $\beta_0 = 0$ and $\beta_1 = 1$ are valid (Froot, 1989). This test has similar implications under the null than the test of the significance of the bias incurred in using forward rates as forecasts presented in table 3. The restrictions are tested using a Wald statistic with variance matrix consistent with heteroscedasticity and autocorrelation using the Newey-West estimator, because of the MA$(h-1)$ in the residuals and the type of data. The interpretation of the coefficients of this regression described in Söderlink and Svensson (1997) implies that if the null with two restrictions is rejected but $\beta_1 = 1$, either there is a constant risk premium or the standard deviation of the risk premium is equal to the negative of the correlation between $E(y_{t+h}^{(3)})$ and the risk premium. This latter combination of parameters is an unlikely cause of not rejecting the null in our case.

The results presented in Table 7 indicate that the restrictions implied by the rational expectations hypothesis ($\beta_0 = 0$ and $\beta_1 = 1$) are in general rejected at 10 per cent level. For the horizons and data periods that we cannot reject the null, the p-values are in general smaller than 15 per cent except for $h = 1$ with the longest data set. However, the t-statistics for testing $\beta_1 = 1$ cannot reject the null with "data 1", indicating a constant risk premium. With the shorter data, there is some evidence that the null of $\beta_1 = 1$ is rejected at intermediary horizons.

This is similar to the summary of the results for the US reported in Söderlink and Svensson (1997).

These results suggest that there are forward premia for all horizons, but there is limited evidence that the forward premia are time-varying. Therefore, there is weak evidence on the rejection of the expectations hypothesis.

Even thought the expectations hypothesis may be valid, it is worthwhile to compute the forward premium because it can help reducing the bias in using the forward rate to forecast the 3-month interest rate. In addition, our procedure of computing confidence intervals (section 3.3) for the forward premium may give us extra information on whether changes in the forward premium are statistically different over time. Finally, because the weak evidence in favour of a time-varying forward premium depends on the sample period, the time-varying behaviour may become more important in the future.

5 The Forward Premium of Euro Interest Rates

In the last section, we present results that support the existence of forward premia, although it may not be time-varying. In the first part of the present section, we present our estimates of the average forward premia for different horizons. In the second part, we evaluate the robustness of the estimates of the forward premium in a real-time exercise to verify the effect of adding information. In the last part, we analyse our estimates along the sample period. We first compare the estimates with its confidence intervals. Finally, we analyse our estimates in comparison with other results in the literature and with measures of market volatility, monetary policy and skewness of the option-implied probability distribution of interest rates futures.

5.1 Forward Premia for Different Horizons

The 3-month forward premia for each horizon are calculated as the difference between the implied 3-month forward rates and the forecasts with the extended Diebold and Li (2006) approach of the 3-month interest rate:

$$\widehat{frp}_t^{(h,h+3)} = \hat{f}_t^{(h,h+3)} - \hat{y}_{t+h|t}^{(3)}.$$

Figure 1 presents the mean of the estimated 3-month forward premia for horizons from 3 to 36 months with and without the inclusion of macroeconomic variables. One can see that the forward premium monotonically increases with the horizon. The fact that the risk premia increase with maturity is a standard result also obtained for German and euro area data in Hordähl et al (2006) and Capiello et al (2006).

The inclusion of macroeconomic variables increases the average of the estimated forward premia. This difference is larger with “data 2”. In general, as discussed in the evaluation of Tables 5 and 6, the reductions in the positive bias are marginal. The inclusion of these additional factors leads also to reductions of the variance of the shocks employed in the computation of the confidence intervals for the forward premia over time. Therefore, we will present estimates of the forward premia only with the inclusion of macroeconomic variables in the remaining of the paper. Our results do not change qualitatively if these variables are removed from the VAR.

5.2 Robustness of the Estimates of the Forward Premia to Different Sample Periods

The advantage of "data 1" is that we have more data points and changes in economic conditions, hence we can evaluate how our procedure for computing the forward premia and their confidence intervals behaves using only information up to t to estimate the VAR. At each point in time, we compute the forward premium $frp_{t|t}^{(n,n+\tau)}$ and its confidence interval $\left[\begin{array}{cc} lo_{t|t}, & up_{t|t} \end{array}\right]$ by bootstrapping its empirical distribution. Previous analyses on the effect of the introduction of the euro on the German term-structure of interest rates have been evaluated by splitting the sample in 1999 (Capiello et al, 2006). Because our modelling approach is simpler to estimate than the dynamic term-structure models, it is easier to use real-time estimation to verify changes in the estimated forward premia and underlying uncertainty.

We use two ways of adding information at each point in time: recursive and rolling. In the recursive estimation, we start with 77 months of observations (6 1/2 years) and at each change of origin, we add one more observation for the estimation of the VAR (eq. 6). The choice of initial number of observations is based on the number of observations of our shorter set of data (from 2000:2-2006:6) in order to capture instabilities from short-sample availability. In the rolling estimation, at each change of origin we move the window of data, implying that the sample size is fixed in 77 observations. The second method is known to be able to smooth the effect of structural breaks.

Figure 2 presents three plots with the forward premium one-year-ahead ($h = 12$). The first plot shows full sample, recursive and rolling estimates of the forward premium. The rolling estimates are comparatively more volatile, and the recursive estimates are highly correlated with the full sample estimates. The large difference between the behaviour of the estimates are in the period between 2000 and 2001: full sample and recursive estimates suggest that the forward premium is small and negative decreasing in 2000 and increasing in 2001, while rolling estimates suggest that it is large, positive, and with the opposite evolution. This is a first evidence that changes in the dynamic correlations of the yield factors may have changed with the introduction of the euro. Note that similar to the results of Hordähl et al (2006), the average risk premia before and after 1999 does not change significantly.

The second plot shows rolling and recursive estimates together with 90 per cent confidences intervals computed by 1000 bootstrap replications. The bootstrap procedure aims at building the predictive density of the forecasted yields and it is described in section 3.3. The confidence intervals computed with rolling samples are generally narrower than with recursive samples. Therefore, the use of all available data up to t increases the estimation uncertainty. The rolling intervals indicate that the estimates of the forward premium are statistically positive in the period between 2000 and the beginning of 2001, but the recursive and full sample estimates suggest a negative premium. Another interesting information obtained by the range of the interval computed with rolling samples is that the variance of the forward premium has decreased. The last value of the range is 2/3 of the initial value. The reduction in the variance of the yield premia after the introduction of the euro has been also detected by Cappiello et al (2006) by breaking the sample into two. The authors argue that the decreasing variance is associated with decreasing incidence of large shocks after 1999. The advantage of our approach is that it endogenously finds the date of the break.

Finally, the third plot presents the forward premium with data 2 and 3, and with "data 1" both with full sample and with only the sample after 1999. The results show that the premium estimated with "data 1" using data only after 1999 are closer to the "data 2" and "data 3" estimates than to the full sample estimates. As a consequence, the disagreement over the forward premium in the 2000-2001 period is caused by the sample period and not by the type of zero-coupon interest rates used.

Therefore, these results suggest a break in the dynamical correlation between the yield factors, conditional on macroeconomic variables, in the period 2000-2001 period. When these changes are not considered, the premium in the years after 1999 is undervalued. The structural break detected could be due to the novelty of the euro as a currency. Thus it may be a rare event and it is unlikely that will occur again. Therefore, our preferred measures of forward premia are based on estimates with data after 1999, using data sets 2 and 3.

5.3 Comparative Evaluation of Estimates of Forward Premia

Figure 3 presents the estimated 3-month forward premia with data 2 and 3 for forecasting horizons 3, 12, 24 and 36 and their 90 per cent confidence intervals. The forward premia are not significantly different from zero for the 3-month and 12-month horizons, even though the point estimates are positive. For longer forecasting horizons (24 and 36 months), the forward premia are significantly positive from 1999:11 up to 2002:6. The variability of the forward premia increases with the horizon because the variability of the forecast yields with the extended Diebold and Li (2006) approach decreases with the horizon. The finding that the variability increases with the horizon is also obtained by Hordähl et al (2006) with an affine term-structure model augmented with a dynamic structural macro model.

The time-varying behaviour of the estimated forward premia has some resemblance with the yield premia computed by Werner (2006) with an affine term-structure model and data after 1995. For the longest horizons, the premia increase in 1999, start to decrease in 2000, and only reverse the downward trend in mid-2005. Our results are also similar to the one-year yield premium estimated with data after 1999 by Capiello et al (2006). The fact that our measure of forward premium does not differ in large scale from the ones presented in the literature based on affine term-structure models gives support to the use of our method, which is less demanding in computation.

To improve our understanding on the time-varying behaviour of our estimates, we compare our estimates of the forward premium with a measure of uncertainty among investors about the future bond yield developments: the implied volatility in the option prices over futures contracts on German bonds. The first plot in figure 4 presents the market volatility together with the forward premium for the 12-month horizon with the three sets of data. There is a positive correlation between the volatility and the forward premium estimated with the full sample for “data 1”. However, in the period 2000-2001, the large values of the estimated forward premia with “data 2” and “data 3” are not associated with high volatility.

The second plot shows that the forward premia estimated with “data 2” and “data 3” are positively correlated with the key ECB interest rate. The association with the official interest rate is not found in the period before the introduction of the euro. Capiello et al (2006) also indicate differences in the correlation between the one-year yield premium and the level of the short-term interest rate: it is negative before the euro and positive after.

An economic interpretation of the relation between the risk premium and short-term interest rates may be the existence of a relation between movements in the official interest rates and the probability attached by market participants to increases in interest rates in the future in comparison to their actual expected values (Vähämaa, 2004). In particular it is natural to expect that investors will demand a higher protection for potential capital looses if there is an increase in the probability that future interest rates will turn out to be higher than their expected values.

One way of assessing this claim is to look at the skewness of the distribution of expected future short-term interest rates. This can be done by using options on EURIBOR futures contracts. We calculate the skewness of option-implied probability distribution of the one-year-ahead 3-month EURIBOR futures contracts for the period after 1999. The option-implied probability distribution is calculated using the one-year-ahead 3-month EURIBOR nearest contract (for example, in January 1999 we use the contract for December 2000, and in February, March and April 1999 the contracts for March 2000). The skewness is measured by the Fisher coefficient, that is, the ratio of the third central moment to the cubed standard deviation. It has a positive (negative) value when the distribution is positively (negatively) skewed, that is, when there is a higher probability that the values stay below (above) the average of the distribution than above (below). The third plot of Figure 4 shows that the skewness is negatively correlated with the 12-month ahead 3-month forward premia estimated with “data 2” and “data 3”. When the skewness is decreasing, the probability distribution is moving to the right, implying that the probability associated with future rises in interest rates as compared to their expected values is increasing. It is important to notice that the futures contracts used in the estimation of the distribution might also incorporate a risk premium. Therefore, increases in the probability attached to future rises in the interest rates may result either from a review of the risk neutral assessment of the likelihood of future rises or from an increase in the risk aversion.

6 Concluding Remarks

We present evidence that forward rates are biased forecasts of euro interest rates. A forward premium explains this bias at short and medium forecast horizons, while unexpected shocks give also a small contribution at long horizons. Using tests with predictive regressions, there is only weak evidence that the premium is time-varying. Nevertheless, the estimates of the forward premium and their statistical significance vary over time. Our estimates also show that the level of the premium and its variability increase with the forecast horizon.

When using German data before the introduction of the euro, we identify a break in the dynamic correlation between yield factors (slope, curvature and level) in the period 2000-2001. Because this break affects the estimates of the forward premium, we advice to use only information after 1999 to compute the euro forward premium.

Even though the forward premia of horizons from 3 to 36 months are on average positive, the observation of the evolution of the forward premia over time together with their confidence intervals indicate that they are significantly equal to zero in some periods of time. The forward premia for the period after 1999 are positively correlated with the key ECB interest rate. One reason for this positive relation might be the fact that when monetary policy is tightening, the market perception changes towards attributing a higher probability to the "bad news" scenario that interest rates could be higher than expected.

In addition to these empirical results, this paper contributes with a competitive method to compute the forward premium and its confidence interval. These are required to evaluate whether a change in the forward rate is due either to modifications in the compensation for risk or in the market expectations. Our approach is easy to estimate and it is flexible to include information from other economically relevant factors in addition to the yield curve factors. It does not impose no-arbitrage restrictions but it is able to capture the time-varying behaviour of the forward premium such as the dynamic term structure models by Capiello et al (2006) and Werner (2006). Another advantage is that the premium and its confidence interval can be computed daily.

The method we have employed for computing the forward premium could also be applied, after some modifications, to understand the recent long-term interest rate "conundrum" (Rudebusch et al, 2006). An explanation usually given for the fact that long-term yields have been at unusually low levels in the US and also in Europe after 2004 is the reduction of the risk premium. The computation of the term premium of long yields with a method similar to the one successfully employed in this paper for the forward premium could help us to evaluate if the reduced risk premium explanation is supported by the data.

A Equivalences of the different formulations of the expectations hypothesis and the measurement of risk premia

The equivalence between the different formulations of the expectations hypothesis

We will show that, in a discrete-time formulation with the variables defined in logs, the four forms of the expectations hypothesis are mathematically equal to each other. Given the relation between the yield and the return $r_t^{(n)} = ny_t^{(n)}$, the RTM written in terms of yields

$$r_t^{(n)} = E_t(r_t^{(1)} + r_{t+1}^{(1)} + ... + r_{t+n-1}^{(1)}) \tag{8}$$

corresponds to the YTM

$$y_t^{(n)} = \frac{1}{n} E_t(y_t^{(1)} + y_{t+1}^{(1)} + ... + y_{t+n-1}^{(1)}). \tag{9}$$

In the case of the UEH, $f_t^{(n,n+1)} = E_t(y_{t+n}^{(1)})$. If we replace the forward rate by its expression in terms of prices, $f_t^{(n,n+1)} = p_t^{(n)} - p_t^{(n+1)}$, and rewrite the UEH with $p_t^{(n+1)}$on the left-hand side, we have:

$$p_t^{(n+1)} = p_t^{(n)} - E_t(y_{t+n}^{(1)}). \tag{10}$$

Solving the previous equation recursively leads to

$$p_t^{(n+1)} = p_t^{(0)} - E_t(y_t^{(1)} - ... - y_{t+n}^{(1)}). \tag{11}$$

Taking into consideration that $p_t^{(0)}$is by definition zero (since $P_t^{(0)} = 1$) and replacing $p_t^{(n+1)}$ by $-(n+1)y_t^{(n+1)}$, we obtain that

$$y_t^{(n+1)} = \frac{1}{n+1} E_t(y_t^{(1)} + y_{t+1}^{(1)} + ... + y_{t+n}^{(1)}),$$

which when it is written for a yield with maturity n is the same as the YTM

For the LEH, $E_t(hpr_{t+1}^{(n)}) = y_t^{(1)}$, if we write it with the $p_t^{(n)}$ on the LEH

$$p_t^{(n)} = E_t(p_{t+1}^{(n-1)}) - y_t^{(1)}, \tag{12}$$

and then solve it recursively, we have

$$p_t^{(n)} = E_t(p_{t+n}^{(0)} - y_{t+n-1}^{(1)} - ... - y_{t+1}^{(1)}) - y_t^{(1)}, \tag{13}$$

which is the same as YTM, given that $p_{t+n}^{(0)}$ is zero and $p_t^{(n)} = -ny_t^{(n)}$.

Relation between the risk premia derived from the different formulations of the expectations hypothesis

Using the definitions in the second section, the following equations define the premia associated with YTM, RTM, UEH and LEH, which we call, respectively, yield premium (yrp_t^n), return premium ($rrp_t^{(n)}$), one-period forward premium n-periods ahead ($frp_t^{(n,n+1)}$) and one-period holding premium ($hrp_t^{(n)}$):

$$yrp_t^{(n)} = y_t^{(n)} - \frac{1}{n}E_t(y_t^{(1)} + y_{t+1}^{(1)} + ... + y_{t+n-1}^{(1)}); \tag{14}$$

$$rrp_t^{(n)} = r_t^{(n)} - E_t(r_t^{(1)} + r_{t+1}^{(1)} + ... + r_{t+n-1}^{(1)}); \tag{15}$$

$$frp_t^{(n,n+1)} = f_t^{(n,n+1)} - E_t(y_{t+n}^{(1)}); \tag{16}$$

$$hrp_t^{(n)} = E_t(hpr_{t+1}^{(n)}) - y_t^{(1)}. \tag{17}$$

To demonstrate the equivalence between the yield premium (14) and the return premium (15), we replace the yields by its expression in terms of returns in the equation 14,

$$yrp_t^{(n)} = \frac{r_t^{(n)}}{n} - \frac{1}{n}E_t(r_t^{(1)} + r_{t+1}^{(1)} + ... + r_{t+n-1}^{(1)}).$$

Thus the yield premium of a zero-coupon bond with maturity n is equal to the return premium of that bond divided by n:

$$yrp_t^{(n)} = \frac{rrp_t^{(n)}}{n}. \tag{18}$$

Next, we show that the yield premium of a zero-coupon bond with maturity n is equal to the average of the one-period forward premia between the present period and $n-1$ periods ahead:

$$yrp_t^{(n)} = \frac{1}{n}\sum_{i=0}^{n-1} frp_t^{(i,i+1)}. \tag{19}$$

This is so because the average of the one-period forward premia can be written as the difference between the average forward rates and the average of expected future one-period yields:

$$\begin{aligned} \frac{1}{n}\sum_{i=0}^{n-1} frp_t^{(i,i+1)} &= \frac{1}{n}\sum_{i=0}^{n-1}\left[f_t^{(i,i+1)} - E_t(y_{t+i}^{(1)})\right] \\ &= \frac{1}{n}\sum_{i=0}^{n-1} f_t^{(i,i+1)} - \frac{1}{n}\sum_{i=0}^{n-1} E_t(y_{t+i}^{(1)}), \end{aligned} \tag{20}$$

where the last term is equal to the last term in the equation for the yield premium (14), and the first term in the RHS corresponds to the average of the one-period forward rates between the present period and $n-1$ periods ahead:

$$\begin{aligned} \frac{1}{n}\sum_{i=0}^{n-1} f_t^{(i,i+1)} &= \frac{1}{n}\left(f_t^{(0,1)} + f_t^{(1,2)} + ... + f_t^{(n-2,n-1)} + f_t^{(n-1,n)}\right) \\ &= \frac{1}{n}\left[\left(p_t^{(0)} - p_t^{(1)}\right) + \left(p_t^{(1)} - p_t^{(2)}\right) + ... + \left(p_t^{(n-2)} - p_t^{(n-1)}\right) + \left(p_t^{(n-1)} - p_t^{(n)}\right)\right] \\ &= -\frac{1}{n}p_t^{(n)} = y_t^{(n)}, \end{aligned} \tag{21}$$

which is equal to the yield of a bond with maturity n.

In similar way, we can show that the yield premium of a zero-coupon bond with maturity n is equal to the average of the one-period holding premia (for that bond) between the present moment and $n-1$ periods ahead:

$$yrp_t^{(n)} = \frac{1}{n}\sum_{i=0}^{n-1} hrp_{t+i}^{(n-i)}. \tag{22}$$

This is a intuitive result since the RHS can be decomposed as

$$\frac{1}{n}\sum_{i=0}^{n-1} hrp_{t+i}^{(n-i)} = \frac{1}{n}\sum_{i=0}^{n-1} E_t(hpr_{t+1+i}^{(n-i)}) - \frac{1}{n}\sum_{i=0}^{n-1} E_t(y_{t+i}^{(1)}), \tag{23}$$

where as before the last term is equal to the last term in the equation for the yield premium (14). The first term on the RHS of (23) is the sum of the expected one-period holding return over the maturity of a bond divided by n that corresponds to its return to maturity divided n, which is the yield to maturity as follows:

$$\begin{aligned} \frac{1}{n}\sum_{i=0}^{n-1} E_t(hpr_{t+1+i}^{(n-i)}) &= \frac{1}{n}\sum_{i=0}^{n-1} E_t(p_{t+i+1}^{(n-i-1)} - p_{t+i}^{(n-i)}) \\ &= \frac{1}{n}[E_t(p_{t+1}^{(n-1)} - p_t^{(n)}) + E_t(p_{t+2}^{(n-2)} - p_{t+1}^{(n-1)}) + ... \\ &\quad + E_t(p_{t+n-1}^{(1)} - p_{t+n-2}^{(2)}) + E_t(p_{t+n}^{(0)} - p_{t+n-1}^{(1)})] \\ &= -\frac{1}{n}p_t^{(n)} = y_t^{(n)}. \end{aligned} \tag{24}$$

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Table 1: Description of the Data.

	Period	Source and Description
Data 1	1986:1-2006:6 (end-of-month)	1 - Maturities 1,3,6,9 and 12 months are German LIBOR rates (EURIBOR after 1999) transformed into continuously compounded rates. 2- Maturities 2 up to 10-years are yields on zero-coupon bonds calculated by the Deutsche Bundesbank derived from observed yields of a group of coupon bonds (listed Federal bonds (Bunds), five-year Federal notes (Bobls) and Federal Treasury notes (Schatze)). For more details, see Deutsche Bundesbank (1997).
Data 2	1999:1-2006:6 (end-of-month)	1- Maturities 1,3,6,9 and 12 months are EURIBOR rates transformed into continuously compounded rates. 2- Maturities 2 up to 10-years are euro Swap rates from Thomson Financial DataStream.
Data 3	2000:2-2006:6 (end-of-month)	Maturities 1,3,6,9,12,15,18,…,117 and 120 months are estimates of zero-coupon rates retrieved from Thomson Financial DataStream (Intercapital Brokers). The estimates are obtained with data from money market, futures and swap interest rates of euro.
Inflation	1986:1-2006:6	Annual growth rate of consumer prices. Until 1998, it is the seasonally adjusted German Index of Consumer Prices from the Bundesbank. From 1999 onwards, it is the seasonally adjusted euro area Harmonized Index of Consumer Prices from Eurostat.
Real activity growth	1986:1-2006:5	Annual growth rate of industrial production excluding construction. Until 1998, it is the seasonally adjusted series for Germany from the Bundesbank. From 1999 onwards, it is the seasonally adjusted euro area series from Eurostat.

Table 2: Comparing Nelson and Siegel (NS) and Svensson (SV) in fitting the yield curve

The entries are the square root of the mean of squared residuals of each maturity. SV is an estimated yield curve with four factors and NS with three factors. NSf has the coefficient that controls which maturity the curvature has a larger weight (θ_{1t}) fixed and equal to 0.0542. The maturity is measured in months. See the data description in Table 1.

	SV			NS			NSf		
τ	Data 1	Data 2	Data 3	Data 1	Data 2	Data 3	Data 1	Data 2	Data 3
1	0.025	0.024	0.018	0.052	0.042	0.059	0.085	0.061	0.073
3	0.037	0.038	0.023	0.043	0.037	0.029	0.048	0.039	0.035
6	0.025	0.020	0.013	0.035	0.025	0.024	0.050	0.033	0.025
9	0.020	0.013	0.011	0.034	0.024	0.040	0.066	0.046	0.046
12	0.029	0.028	0.015	0.039	0.031	0.044	0.078	0.052	0.051
24	0.030	0.025	0.014	0.050	0.053	0.027	0.051	0.053	0.031
36	0.010	0.010	0.008	0.022	0.022	0.022	0.038	0.032	0.022
48	0.013	0.012	0.006	0.022	0.022	0.018	0.042	0.030	0.023
60	0.008	0.011	0.007	0.025	0.034	0.022	0.040	0.036	0.030
72	0.005	0.007	0.004	0.022	0.029	0.021	0.032	0.031	0.028
84	0.009	0.010	0.008	0.014	0.015	0.011	0.018	0.017	0.015
96	0.009	0.012	0.009	0.008	0.008	0.007	0.011	0.009	0.008
108	0.004	0.005	0.004	0.017	0.019	0.019	0.027	0.022	0.024
120	0.010	0.013	0.011	0.031	0.031	0.030	0.048	0.036	0.042

Table 3: Forecast biases with the forward rates

The entries are the mean of forecast errors. The forecast errors are computed using the observed 3-month interest rate ($y^{(3)}_{t+h}$) and the implied forward rate for the 3-month rate computed at *t*. The 3-month forward rate with the maturity in months is $\hat{f}_t^{(n,n+3)} = \frac{1}{3}\left[(n+3)\hat{y}_t^{(n+3)} - (n)\hat{y}_t^{(n)}\right]$. The forward is computed using fitted yields with Nelson and Siegel (NS) with fixed θ_{1t} and Svensson (SV). For example, for t+18, we have $\hat{f}_t^{(18,18+3)}$. The values in brackets are standard deviations of the forecast errors computed using the Newey-West estimator with lag truncation *h-1*. The emboldened values have t-statistics larger than two, implying that the null that the forecast error is equal to zero is rejected. See the data description in Table 1.

	Data 1 (86:1-06:6)		Data 2 (99:1-06:6)		Data 3 (00:2-06:6)	
	SV	NSf	SV	NSf	SV	NSf
t+3	0.081	0.064	0.097	**0.131**	0.089	0.111
	(0.044)	(0.046)	(0.057)	(0.061)	(0.066)	(0.071)
t+6	0.176	0.171	0.265	0.278	**0.334**	**0.340**
	(0.091)	(0.098)	(0.138)	(0.149)	(0.117)	(0.129)
t+9	0.296	0.302	0.478	0.457	**0.666**	**0.651**
	(0.162)	(0.167)	(0.253)	(0.262)	(0.142)	(0.150)
t+12	0.436	0.448	0.722	0.673	**1.004**	**0.972**
	(0.243)	(0.242)	(0.364)	(0.365)	(0.147)	(0.151)
t+15	0.578	0.592	**0.966**	**0.900**	**1.332**	**1.291**
	(0.323)	(0.314)	(0.448)	(0.441)	(0.143)	(0.143)
t+18	0.722	0.735	**1.234**	**1.164**	**1.637**	**1.594**
	(0.400)	(0.385)	(0.480)	(0.464)	(0.147)	(0.146)
t+21	0.871	0.882	**1.514**	**1.450**	**1.887**	**1.849**
	(0.475)	(0.455)	(0.472)	(0.450)	(0.160)	(0.162)
t+24	1.013	1.021	**1.791**	**1.741**	**2.097**	**2.069**
	(0.547)	(0.526)	(0.439)	(0.414)	(0.176)	(0.183)
t+27	1.144	1.150	**2.036**	**2.004**	**2.303**	**2.289**
	(0.616)	(0.595)	(0.397)	(0.374)	(0.189)	(0.204)
t+30	1.272	1.276	**2.250**	**2.240**	**2.487**	**2.487**
	(0.679)	(0.661)	(0.345)	(0.327)	(0.218)	(0.240)
t+33	1.408	1.411	**2.414**	**2.426**	**2.633**	**2.649**
	(0.732)	(0.717)	(0.307)	(0.297)	(0.264)	(0.290)
t+36	1.532	**1.534**	**2.516**	**2.548**	**2.721**	**2.751**
	(0.776)	(0.765)	(0.296)	(0.294)	0.313)	(0.336)

Table 4: Granger-Non-Causality tests between yield factors and macroeconomic variables.

The entries are t-statistics for testing whether the variables indicated in the columns do not cause the variables in the rows. The statistics are based on the estimation of a VAR with order 1 with the three yield factors and the two macroeconomic variables as endogenous. The indication (-) in the slope means that we take the estimate directly from the Nelson and Siegel (eq. 3), so it has opposite sign to the traditional measure of slope. See the data description in Table 1.

	Growth causes …	Inflation causes …	Growth is caused by …	Inflation is caused by …
Data 1				
Level	0.161	-1.355	-0.567	0.425
Slope(-)	0.877	**1.957**	**-3.704**	1.058
Curvature	**2.548**	0.561	**4.109**	0.599
Data 2				
Level	-1.416	**-4.110**	-0.175	0.203
Slope(-)	**2.577**	**3.252**	-1.431	-0.053
Curvature	0.108	-0.977	**3.019**	0.562
Data 3				
Level	-1.477	-0.419	-0.267	0.025
Slope (-)	**2.707**	-0.054	-0.605	**1.992**
Curvature	-0.542	-0.663	**2.159**	-0.002

Table 5: Forecast biases with the extended Diebold and Li (2006) approach

The entries are the mean of forecast errors. The forecast errors are computed using the observed three-month interest rates ($y^{(3)}_{t+h}$) and the forecasts of the 3-month rate obtained using the approach described in section 3. The VAR employed in the forecast includes in addition to the yield factors the inflation and economic growth when indicated (w. Macro). The values in brackets are standard deviations of the forecast errors computed using the Newey-West estimator with lag truncation *h-1*, where *h* is the forecast horizon. The emboldened values have t-statistics larger than two, implying that the null that the forecast error is equal to zero is rejected. See the data description in Table 1.

	Data 1 (86:1-06:6)		Data 2 (99:1-06:6)		Data 3 (00:2-06:6)	
		w. Macro		w. Macro		w. Macro
T+3	0.003 (0.046)	0.001 (0.045)	-0.008 (0.062)	-0.011 (0.056)	0.038 (0.072)	0.026 (0.063)
T+6	0.008 (0.099)	0.004 (0.097)	-0.045 (0.137)	-0.046 (0.105)	0.081 (0.129)	0.052 (0.113)
t+9	0.015 (0.169)	0.010 (0.166)	-0.061 (0.237)	-0.064 (0.171)	0.164 (0.159)	0.121 (0.150)
t+12	0.025 (0.245)	0.020 (0.242)	-0.040 (0.328)	-0.056 (0.231)	0.242) (0.173)	0.191 (0.174)
t+15	0.025 (0.319)	0.020 (0.316)	0.003 (0.396)	-0.040 (0.279)	0.322 (0.173)	0.268 (0.184)
t+18	0.022 (0.391)	0.017 (0.386)	0.094 (0.419)	0.016 (0.303)	**0.403** (0.169)	0.350 (0.188)
t+21	0.025 (0.463)	0.020 (0.455)	0.223 (0.410)	0.105 (0.310)	**0.458** (0.175)	**0.410** (0.197)
t+24	0.023 (0.537)	0.018 (0.525)	0.372 (0.380)	0.213 (0.302)	**0.505** (0.186)	**0.464** (0.208)
t+27	0.016 (0.612)	0.011 (0.596)	0.508 (0.341)	0.313 (0.286)	**0.578** (0.175)	**0.543** (0.194)
t+30	0.012 (0.684)	0.007 (0.665)	**0.629** (0.284)	0.403 (0.251)	**0.656** (0.138)	**0.625** (0.153)
t+33	0.024 (0.747)	0.018 (0.723)	**0.713** (0.225)	**0.463** (0.208)	**0.718** (0.085)	**0.691** (0.093)
t+36	0.032 (0.802)	0.025 (0.775)	**0.743** (0.180)	**0.476** (0.174)	**0.741** (0.044)	**0.714** (0.048)

Table 6: Forecast biases with the extended Diebold and Li (2006) approach with common sample

Notes as Table 5. It uses estimates for sample 2000:2-2006:6.

	Data 3	Data 1		Data 2		Data 3	
	Using Forwards		with Macro		with Macro		with Macro
t+3	0.111 (0.071)	0.046 (0.064)	0.040 (0.058)	0.031 (0.066)	0.022 (0.059)	0.038 (0.072)	0.026 (0.063)
t+6	**0.340** (0.129)	0.074 (0.113)	0.061 (0.104)	0.069 (0.117)	0.047 (0.106)	0.081 (0.129)	0.052 (0.113)
t+9	**0.651** (0.150)	0.133 (0.144)	0.116 (0.144)	0.140 (0.147)	0.109 (0.143)	0.164 (0.159)	0.121 (0.150)
t+12	**0.972** (0.151)	0.187 (0.157)	0.169 (0.166)	0.205 (0.164)	0.169 (0.168)	0.242 (0.173)	0.191 (0.174)
t+15	**1.291** (0.143)	0.244 (0.157)	0.229 (0.174)	0.271 (0.167)	0.235 (0.180)	0.322 (0.173)	0.268 (0.184)
t+18	**1.594** (0.146)	0.306 (0.154)	0.296 (0.176)	**0.341** (0.163)	0.308 (0.183)	**0.403** (0.169)	0.350 (0.188)
t+21	**1.849** (0.162)	**0.349** (0.161)	0.346 (0.184)	**0.389** (0.169)	0.362 (0.191)	**0.458** (0.175)	**0.410** (0.197)
t+24	**2.069** (0.183)	**0.388** (0.175)	**0.390** (0.194)	**0.429** (0.179)	**0.410** (0.200)	**0.505** (0.186)	**0.464** (0.208)
t+27	**2.289** (0.204)	**0.455** (0.169)	**0.461** (0.182)	**0.496** (0.170)	**0.483** (0.188)	**0.578** (0.175)	**0.543** (0.194)
t+30	**2.487** (0.240)	**0.527** (0.139)	**0.535** (0.143)	**0.566** (0.137)	**0.559** (0.149)	**0.656** (0.138)	**0.625** (0.153)
t+33	**2.649** (0.290)	**0.589** (0.090)	**0.595** (0.087)	**0.625** (0.087)	**0.620** (0.093)	**0.718** (0.085)	**0.691** (0.093)
t+36	**2.751** (0.336)	**0.615** (0.045)	**0.616** (0.044)	**0.647** (0.044)	**0.643** (0.047)	**0.741** (0.044)	**0.714** (0.048)

Table 7: Test of the Expectations Hypothesis

The first entries for each database are estimates (with Newey-West robust standard errors in parenthesis) of the parameters of the equation $(y_{t+h}^{(3)} - y_t^{(3)}) = \beta_0 + \beta_1 (f_t^{h,h+3} - y_t^{(3)}) + \varepsilon_{t+h}$. The last two entries are p-values of the test with the null indicated. The emboldened values are p-values that indicate rejection of the null at 10 per cent significance level. See the data description in Table 1.

	Data 1 (86:1-06:6)				Data 2 (99:1-06:6)				Data 3 (00:2-06:6)			
	β_0	β_1	β_1=1	β_0=0; β_1=1;	β_0	β_1	β_1=1	β_0=0; β_1=1;	β_0	β_1	β_1=1	β_0=0; β_1=1;
t+3	-0.06 (0.04)	0.89 (0.15)	0.11	0.38	-0.17 (0.05)	1.51 (0.26)	**0.01**	**0.00**	-0.10 (0.07)	1.39 (0.36)	0.07	**0.06**
t+6	-0.17 (0.09)	0.95 (0.18)	0.20	0.12	-0.33 (0.12)	1.28 (0.33)	0.10	**0.01**	-0.34 (0.12)	1.17 (0.41)	0.17	**0.00**
t+9	-0.29 (0.14)	0.96 (0.24)	0.22	**0.08**	-0.52 (0.21)	1.19 (0.43)	0.16	**0.05**	-0.64 (0.13)	0.92 (0.43)	0.21	**0.00**
t+12	-0.42 (0.21)	0.90 (0.29)	0.19	0.14	-0.68 (0.29)	1.02 (0.44)	0.24	**0.06**	-0.90 (0.13)	0.63 (0.47)	0.11	**0.00**
t+15	-0.55 (0.28)	0.91 (0.34)	0.20	0.14	-0.77 (0.38)	0.80 (0.38)	0.15	0.12	-1.10 (0.14)	0.36 (0.39)	**0.03**	**0.00**
t+18	-0.71 (0.35)	0.96 (0.37)	0.23	0.12	-0.90 (0.49)	0.65 (0.31)	**0.07**	**0.06**	-1.27 (0.21)	0.26 (0.27)	**0.00**	**0.00**
t+21	-0.88 (0.41)	1.00 (0.38)	0.25	0.10	-1.10 (0.58)	0.60 (0.29)	**0.04**	**0.00**	-1.48 (0.29)	0.34 (0.25)	**0.00**	**0.00**
t+24	-1.07 (0.48)	1.06 (0.39)	0.22	**0.08**	-1.39 (0.64)	0.66 (0.31)	**0.07**	**0.00**	-1.78 (0.36)	0.58 (0.29)	**0.04**	**0.00**
t+27	-1.26 (0.53)	1.12 (0.39)	0.19	**0.06**	-1.66 (0.70)	0.70 (0.33)	**0.09**	**0.00**	-2.14 (0.40)	0.81 (0.28)	0.13	**0.00**
t+30	-1.45 (0.58)	1.18 (0.39)	0.16	**0.04**	-2.03 (0.65)	0.84 (0.30)	0.15	**0.00**	-2.44 (0.42)	0.95 (0.23)	0.21	**0.00**
t+33	-1.65 (0.60)	1.22 (0.39)	0.14	**0.02**	-2.32 (0.57)	0.92 (0.25)	0.19	**0.00**	-2.65 (0.42)	1.00 (0.17)	0.25	**0.00**
t+36	-1.86 (0.61)	1.28 (0.38)	0.12	**0.01**	-2.36 (0.58)	0.87 (0.22)	0.14	**0.00**	-2.79 (0.39)	1.04 (0.14)	0.20	**0.00**

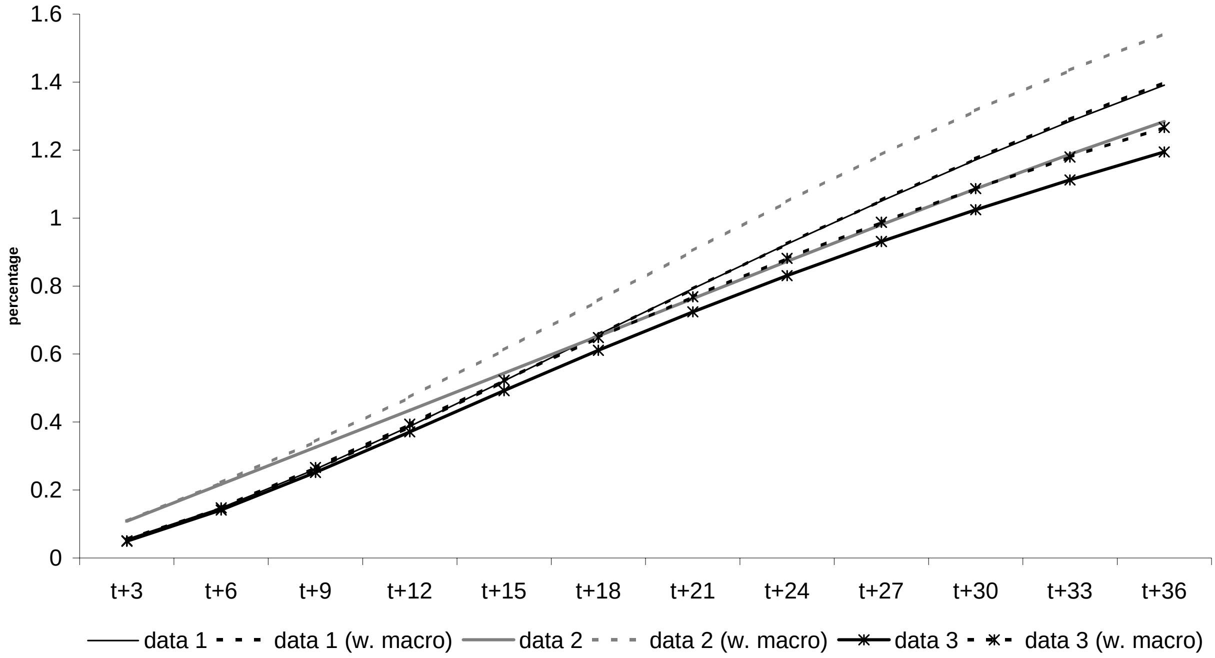


Figure 1: Mean of the 3-month forward premia for each horizon.

The "*w. macro*" means that the VAR employed in the computation of the forward premia includes not only the three factors of the yield curve but also the macro variables. See Table 1 for description of the data.

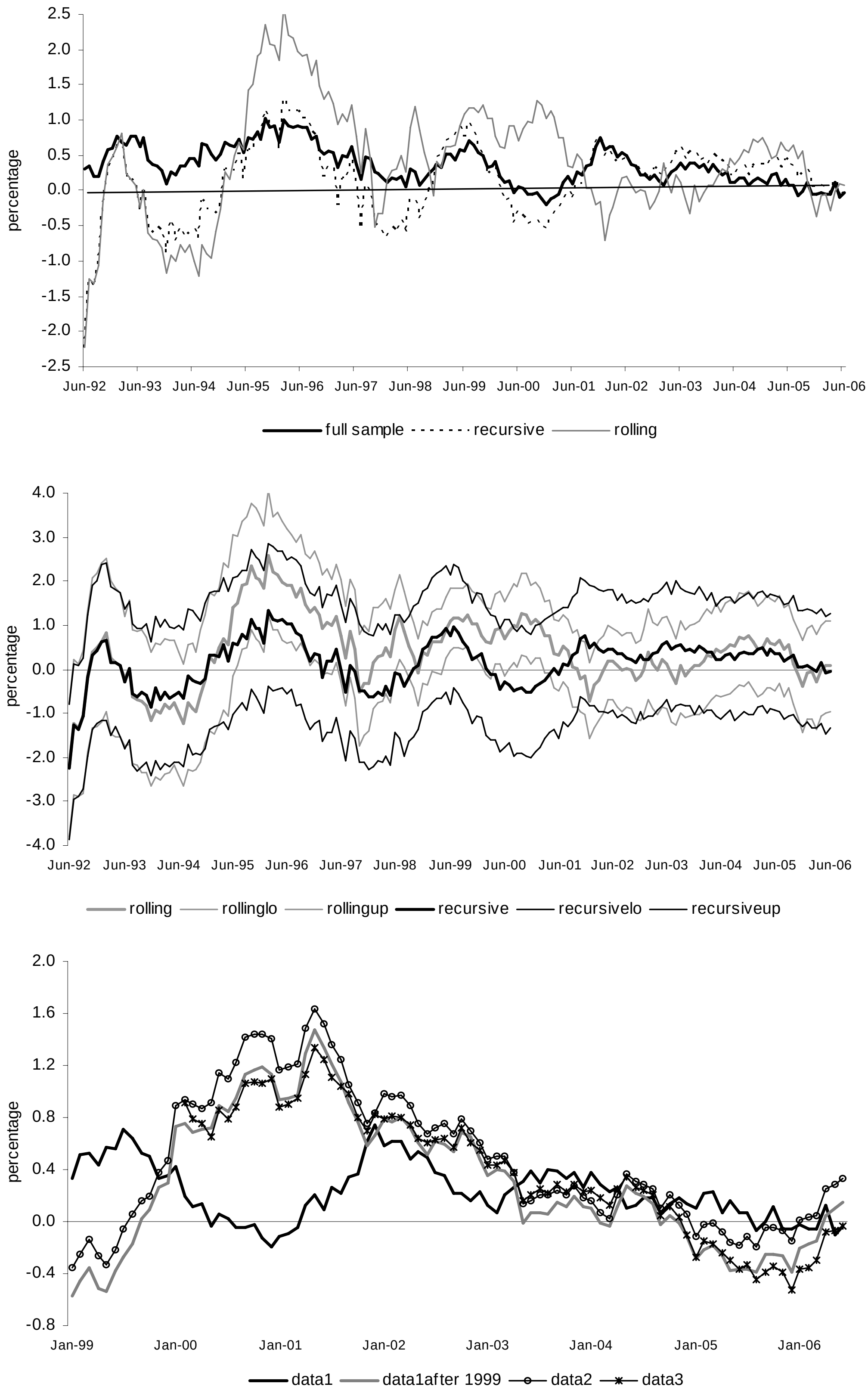


Figure 2: Estimates of the forward premia and 90 per cent confidence intervals for t+12

The estimates are obtained including macro variables in the VAR. The top two charts include estimates computed with “data1”. The first one compares the estimates with the full sample with the recursive and rolling estimates. The rolling windows have a size of 77 months. The second plot includes the recursive and rolling estimates as well the lower (lo) and upper (up) limits of their 90 per cent confidence intervals, which were obtained with 1000 boostraps, as described in section 3.3. The bottom plot includes the full sample estimations for the three sets of data and estimates with “data1” with sample starting in 1999. See Table 1 for description of the data.

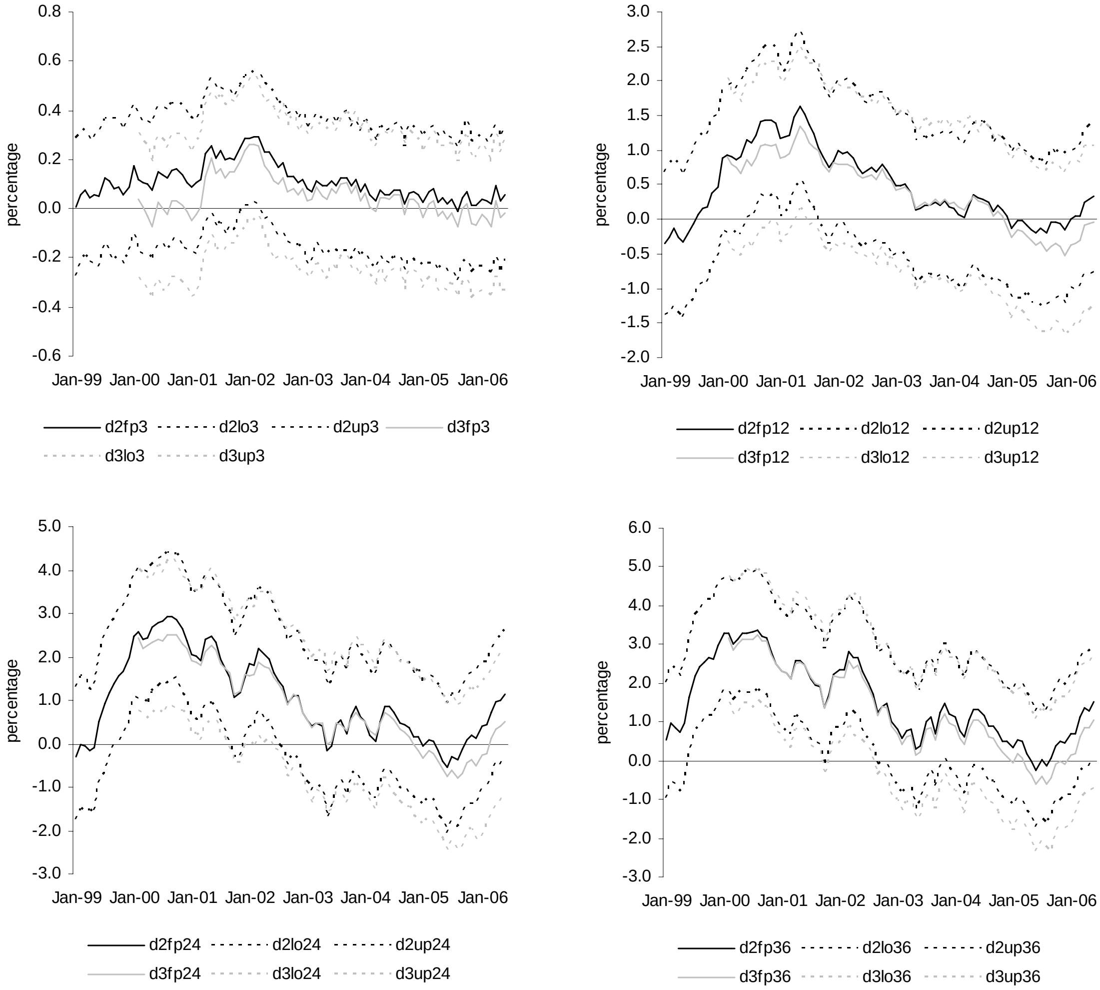
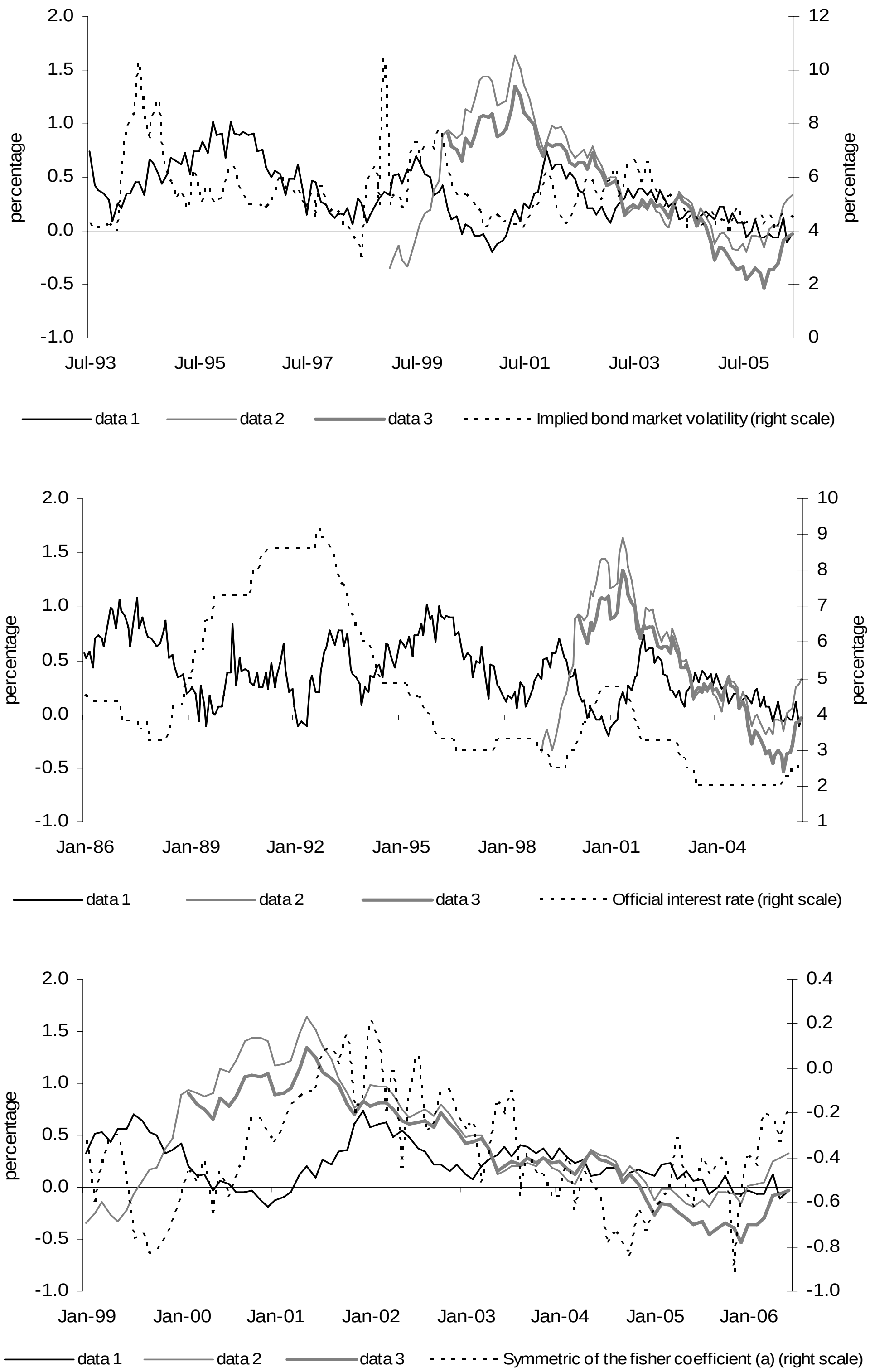


Figure 3: Estimates of the forward premia and 90 per cent bootstrapped confidence intervals at t+3, t+12, t+24, t+36.

Estimates are computed with data 2 (d2) and 3 (d3) (see description in Table 1). The estimates use information of macro variables. The estimated premia is *fp*; the lower limit of the confidence interval is *lo*, and the upper is *up*. 1000 bootstrap replications are employed to compute the confidence intervals.



(a) Fisher coefficient of the option-implied probability distribution of the 3-month EURIBOR futures contracts for t+12.

Figure 4: Forward premia at t+12, volatility, policy rate and skewness of market expectations.

Estimates of the forward premia are obtained with information on macro variables. For description of the data, see Table 1. The implied bond market volatility corresponds to the 5-days average of the end-of-month implied volatility in the options prices over the nearest contract of German Government bond with a maturity of 8.5-10.5 years traded in the Eurex Deutschland (Source: Bloomberg). The official interest rate is the end of month values of the Bundesbank repo rate until December 1998 and of the ECB main refinancing operations interest rate after January 1999. The Fisher coefficient measures the skewness and is positive (negative) when the distribution is positively (negatively) skewed, that is, when there is a higher probability that the values stay below (above) the average of the distribution than above.

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