The term structure as a predictor of real activity and inflation in the euro area: a reassessment

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1. Introduction

The slope of the yield curve is often used by financial and policy analysts as an indicator of future real activity and inflation. Empirical research tends to confirm the predictive power of the yield spread both for real activity and for inflation. Empirical research has focused mostly on the US economy and, to some extent, on the larger individual pre-EMU EU countries, whereas there are hardly any estimates for the euro area.[2]

Berk and van Bergeijk (2000, 2001) attempted an empirical assessment for the euro area. They concluded that both for individual euro area countries and for the euro area as a whole, the yield spread contains only very limited information on future inflation rate and output growth changes beyond the information contained in the history of the latter variables.

The present paper makes a new attempt to evaluate empirically the predictive power of the yield spread for euro area output and inflation. It makes use of the longer time series that have become available since Berk and van Bergeijk (2000, 2001). More importantly, the paper proposes a simple method to estimate time-varying term premia that may have caused the poor forecasting performance of the yield spread quoted in the above-cited contribution. We believe that this issue is of particular relevance when working with longer euro area financial market series, since the pre-1999 part (ie normally the larger part!) of the series is usually composed of raw country aggregates potentially plagued by changing risk premia related, among other factors, to the exchange rate mechanism of the European Monetary System. Equally, convergence phenomena in the run-up to the start of EMU may have heavily influenced national European bond rates. Working with synthetic pre-EMU bond rates for the euro area which are not adjusted for these changing risk premia can be expected to strongly influence empirical estimates of economic relationships.[3] In a recent contribution, Carstensen and Hawellek (2003) show that, for German data, assessing the time-varying nature of the term premium improves the quality of inflation forecasts obtained using term structure models.

In this contribution, we show that using a simple adjustment method for risk premia contained in bond rates significantly improves the information content of the term spread for future euro area output and, to a lesser extent, for future inflation rates. The basic idea behind the adjustment procedure is to approximate the (time-varying) term premium by making use of the relationship implied by the rational expectations hypothesis of the term structure (henceforth REHTS). By means of an out-of-sample forecasting exercise, we provide evidence that, for forecasting horizons ranging up to two years, the yield curve adjusted for risk premium improves significantly upon the observed term spread as a predictor of industrial production in the euro area. The results for the inflation rate are less clear-cut, but indicate that the use of the term premium adjustment can lead to improvements in the accuracy of the forecasts of inflation and core inflation rates.

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[2] For an extensive survey of the literature on using asset prices to forecast growth and inflation, see Stock and Watson (2003).

[3] An example of this influence in the context of the estimation of monetary policy reaction functions is given by Crespo Cuaresma et al (2004).

The remainder of the paper is structured as follows. Section 2 summarises the theory underlying the predictive capabilities of the term spread for output and inflation, including the conditions by which they are influenced and limited. Section 3 proposes a simple risk premium adjustment method for euro area bond rates. Section 4 presents evidence for the euro area on the predictive content of the term spread for real activity and inflation, juxtaposing the results based on the premia-adjusted term spreads against results from unadjusted series. Section 5 concludes.

2. Theoretical underpinnings for a leading indicator property of the term structure

The theoretical background underlying the use of the term structure of interest rates as an indicator for market expectations of future inflation and/or real growth is based on the combination of the Fisher equation and the REHTS. The REHTS states that the yield to maturity of a bond with n periods to maturity can be decomposed into expected one-period yields and a risk premium, so that

$$R(n,t) = \frac{1}{n}\sum_{i=0}^{n-1} E_t R(1,t+i) + \Phi(n,t) \tag{1}$$

where $E_t(\cdot)$ is the conditional expectation operator using the information available up to period t, $R(n,t)$ is the yield to maturity of a bond with n periods to maturity, and $\Phi(n,t)$ is the average risk premium on an n-period bond until it matures.

Using the Fisher decomposition, equation (1) can be rewritten as

$$R(n,t) = E_t r(n,t) + E_t \pi(n,t) + \Phi(n,t) \tag{2}$$

where $E_t r(n,t)$ is the average real ex ante interest rate over the periods t to $t+n-1$, and $E_t\pi(n,t)$ is the average expected inflation rate over the periods $t+1$ to $t+n$. Under the REHTS, the risk premium is assumed to be constant over time. We will address this restrictive assumption in the next section.

The slope of the yield curve between maturities m and n can be decomposed into changes in the real rate and in expected inflation making use of (2). Consider equation (2) for a long-term interest rate of maturity n and a short-term interest rate of maturity m. Subtracting the latter from the former, we obtain

$$R(n,t) - R(m,t) = E_t[r(n,t) - r(m,t)] + E_t[\pi(n,t) - \pi(m,t)] + \Phi(n,t) - \Phi(m,t) \tag{3}$$

If real activity is related to changing real interest rates and if the term premium is constant, then equations (2) and (3) imply that the term spread should contain information about future economic activity and inflation.

While the literature on the theoretical background of the relationship between the term spread and future inflation rates is, to the knowledge of the authors, exclusively based on the Fisher decomposition and the REHTS[4] as described above (see Tzavalis and Wickens (1996)), different theoretical underpinnings have been proposed to the link between the term spread and output growth. From a theoretical point of view, the term spread may be related positively or negatively to future real output, depending on the channel at work. Various explanations have been put forward in the literature (see eg Estrella and Mishkin (1997), Berg and van Bergeijk (2000, 2001) and Estrella (2003)).

A first channel derives from the “common factor” effect of current monetary policy on both the term spread and real activity. As a credible central bank, for instance, tightens monetary policy, short-term interest rates rise, while long-term rates rise by less or are not affected at all, leading to a flattening of a previously positively sloped yield curve. After a lag of a few quarters, real activity is also dampened by the restrictive policy. Given the faster reaction of the term spread, the latter leads the slowdown in economic activity.

A second channel works through expectations about future monetary policy changes, in the presence of nominal rigidities. For instance, the expectation of a future monetary tightening (which can be

[4] Notable exceptions, discussed below, are Smets and Tsatsaronis (1997) and Estrella (2003).

thought of as a future shift of the LM curve) would imply higher future short-term rates, thus higher current long-term rates, and, consequently, an increase in the term spread. The expected upward shift in the future LM curve implies a shift to the left in the current IS curve and a fall in current and future output.

A third channel operates through real demand shocks. In terms of a standard IS/LM framework, an expected economic upswing as represented by a future outward shift in the IS curve raises expected future short-term rates (the expected outward shift in the IS curve raises future money demand). Due to the REHTS arbitrage condition, this expectation translates into higher current long-term rates.

In a fourth category of explanations, Harvey (1988) and Hu (1993) explain the correlation between the term spread and future economic growth from intertemporal consumption smoothing by using the consumption capital asset pricing model. The first-order condition of the consumption-based asset pricing model proposed by Campbell (1988) implies that expected returns and consumption growth are linearly related. Consequently, one should observe a comovement between the (real) term structure and the business cycle.

Two attempts to embed the link between the term spread and real activity into a broader analytical framework warrant specific mentioning. Smets and Tsatsaronis (1997) model the joint movements of output, inflation and the nominal term structure as the combined effect of four distinct fundamental shocks: aggregate demand, aggregate supply, monetary policy, and a long-term interest rate shock (driven by unwarranted “inflation scare”). They find that in both Germany and the United States about half of the medium-term variability in the term spread is accounted for by demand and monetary policy shocks, the other half being driven by supply shocks in Germany but by fears about long-term inflation prospects in the United States. They attribute this difference to the higher anti-inflationary credibility enjoyed by the Deutsche Bundesbank. They also find that the big role of supply shocks in explaining term-spread variability is the main reason for the much stronger leading indicator properties of the term spread for output growth. Finally, they show that the predictive content of the term spread is time-varying.

Estrella (2003) systematically investigates factors influencing the predictive power of the term spread for inflation and real variables in the framework of a single formal model comprising a (backward- or forward-looking) Phillips curve, a (backward- or forward-looking) IS equation, the Fisher equation, the term structure, and various monetary policy reaction functions. He finds that the yield curve should be a useful predictor of output and inflation under most circumstances. A positive relationship between the term spread and future output is predicted by the backward-looking form of the model. The prediction capabilities of the term spread importantly depend on the specific form of the policy reaction function. Thus, the predictive relationship, though robust, is not “structural”. In most specifications, further information beyond the term spread is useful in forecasting output. Finally, he finds that, since 1987, reflecting a regime of “strict inflation targeting”, the predictive power of the yield spread, though not entirely absent, has been diminished.

What are the implications from the theoretical literature for the paper at hand? First, there are sufficient sound theoretical underpinnings to justify a further investigation of the empirical leading indicator properties of the term spread for the euro area. Second, most channels and models suggest a positive relationship between the lagged term spread and real activity. However, there are also channels and shocks suggesting a negative relationship. Third, the monetary policy regime may affect the predictive power of the term spread. Thus, any reading of empirical relationships between the term spread and real activity or inflation requires a structural interpretation against the background of prevailing economic circumstances and the monetary policy regime in place. The regime change implied by the transition from the ERM to EMU appears to deserve particular attention in this context. The remainder of this paper pursues this latter aspect by proposing an empirical method to gauge the time-varying term premium in the euro area in the run-up to EMU. We take the theoretical literature as mere background motivation for our research and do not attempt to assess the validity of any of the above theoretical channels. Instead, we will exclusively concentrate on empirically assessing the out-of-sample forecasting abilities of the term spread for output growth and inflation taking into account the time-varying nature of the risk premium.

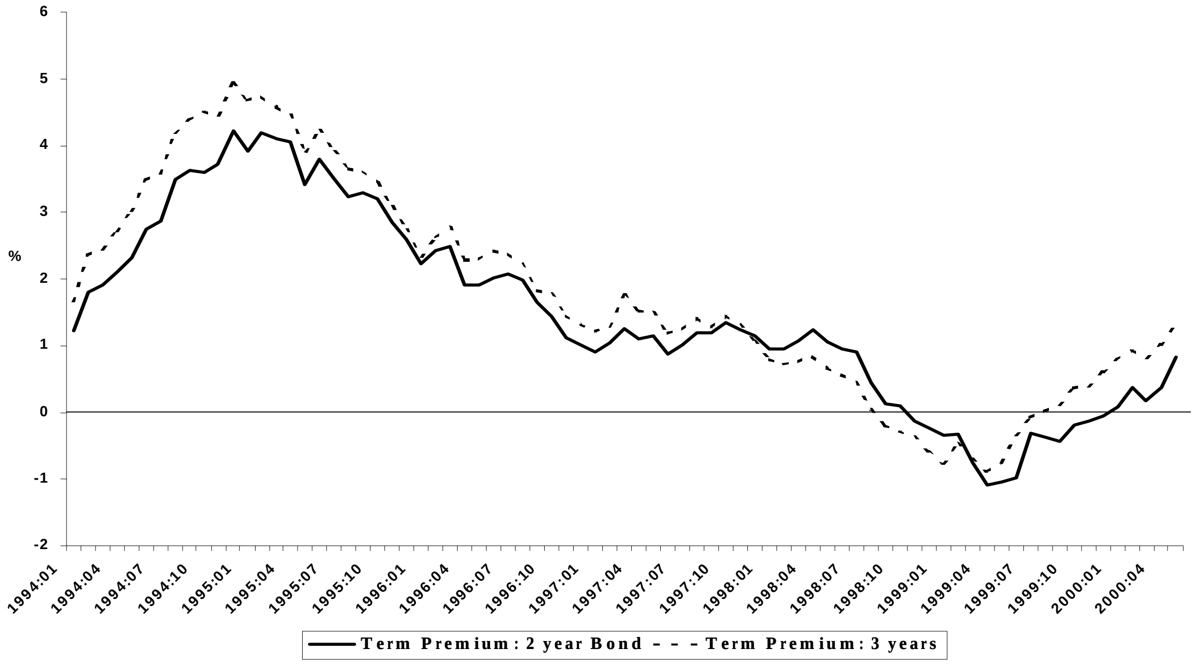
3. A simple risk premium adjustment

The assumption of constant risk premia is unlikely to have held in individual euro area countries and therefore in the euro area as a whole during the time of the ERM and in the run-up to EMU. This section will provide evidence of the existence of time-varying risk premia for long rates in the aggregate euro area. We also propose a simple method to obtain a (potentially time-varying) estimate of $\Phi(n,t)$ in (1).

Some evidence for our claim can be found by extracting the risk premium from equation (1) for the observed two- and three-year bond yields in the euro area.[5] Graph 1 presents the risk premium estimates implied by (1) for these maturities under the assumption of perfect foresight, that is, substituting the expected values with those which were actually realised. The one-month interest rate was used as the short rate. The implied risk premia are plotted for the period ranging from January 1994 (first available observation) to April 2000 (last period for which it is possible to obtain an implied premium for the three-year bond). The risk premium is far from being constant for both cases, reaching a global maximum in late 1994 of around 4 percentage points for the two-year bond and 5 percentage points for the three-year bond. A clear convergence pattern towards zero is observed during the run-up to EMU, culminating in premia around zero in the second half of 1998. The negative risk premium for both long rates during practically the whole of 1999 is due to the increase of short-run nominal interest rates which ran parallel to the rise in inflation after the inception of EMU. The subsequent stabilisation of inflation rates, which was followed by a reduction in the one-month nominal interest rate in the last period of the sample, results in positive risk premia from January 2000 onwards.

Graph 1

Implied risk premia, perfect foresight



However, the estimates presented in Graph 1 can only be obtained a posteriori. If the aim is to correct the term spread for time-varying term premia in order to use the information contained in the adjusted yield curve for predicting future growth rates of output or inflation rates, a real-time estimate of the risk

5 Much of the empirical literature tends to use the 10-year bond on the long side of the term spread. The relatively short sample existing for the aggregate euro area does not allow for sensible empirical work based on such long maturities if the REHTS is to be taken literally.

premium needs to be obtained with information ranging up to the time period in which the forecasts are carried out. We will use a simple expectation formation method to overcome this difficulty. For each time period, we will assume that expectations are formed as forecasts of the variables of interest (the nominal short rate) given the history of this variable up to period t. We will assume that individuals obtain point forecasts of the short-term nominal interest rate using simple autoregressive models. Using the information up to period t on one-month nominal rates, an autoregressive process of order p (AR(p)) model[6] is fitted to the data, and forecasts of the short rate are obtained for $n-1$ periods ahead, where n is the maturity of the bond whose risk premium we are estimating.

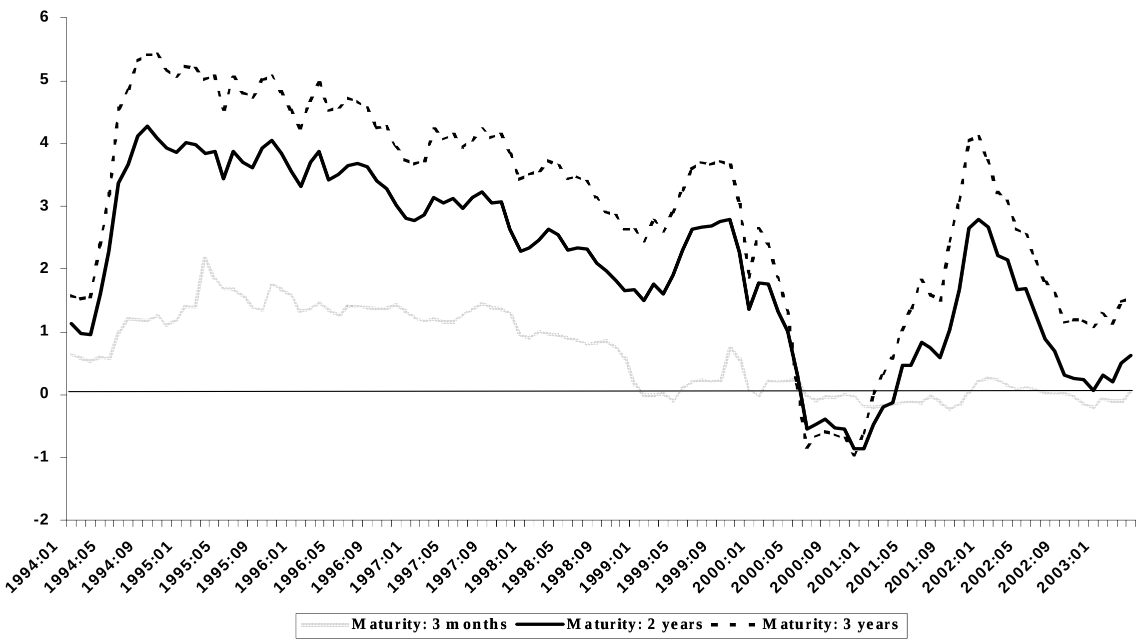
The estimate of the risk premium of the bond with maturity n at period $t(\hat{\Phi}(n,t))$ is then given by the difference between the actual bond yield and the yield implied by the first terms on the right-hand side of (1)

$$\hat{\Phi}(n,t) = R(n,t) - \frac{1}{n}\sum_{i=0}^{n-1}\hat{R}(1,t+i) \tag{4}$$

where $\hat{R}(1,t+i)$ is the one-month real interest rate in period $t+i$ predicted by the autoregressive model. Analogously to the definition of $\Phi(n,t)$ in (1) if perfect foresight is not assumed, the estimate given by (4) is not only composed of a risk premium, but also includes the forecast error of individuals when forming expectations.

Graph 2

Risk premia estimates: autoregressive expectations



Graph 2 presents the estimates of the risk premia obtained by applying this method to the euro area data for three-month interest rates and the two- and three-year bond yields.[7] Significant deviations from zero, ranging up to 200 basis points, appear already for the three-month interest rate in the pre-EMU sample, with a downward-sloping trend since 1996. The term premium associated with the three-month interest rate practically disappears for the EMU period. The overall dynamics and range

[6] A trend was included in the AR(p) specification to account for the departure from stationarity which is observable in the short-term nominal interest rate series for the euro area. At each time period, the length of the AR(p) model was chosen to be the one that minimises the Schwarz criterion among lags one to 12.

[7] See the Appendix for a description of the data and their source.

of the term premium for the three-month interest rate resemble closely the estimates obtained by Crespo Cuaresma et al (2004), who model pre-EMU interest rate spreads with the German short-term interest rate as depending upon expected inflation and output gap differences. For the long-term interest rates, the pre-EMU convergence to a zero term premium occurs with some delay compared to the three-month interest rate and is followed by a resurgence in the risk premia in the EMU period. The risk premia for the long-term rates estimated by this method present more persistence and higher values in the first part of the sample compared to the perfect foresight case due to the fact that the AR(*p*) model produced downward-sloping projections of the short rate also for the period where the one-month interest rate showed a stable dynamic pattern. The same line of reasoning applies to the increase in risk premia after 2001, where the decrease in nominal interest rates observed in the data was expected, according to the projections of the AR(*p*) model, to continue for longer than it actually did.

4. The predictive content of the term spread for real activity and inflation: evidence for the euro area

The results in the previous section suggest that the assumption of a constant risk premium may not hold for euro area data spanning long enough periods of time. This section investigates whether the predictive abilities of the term spread for industrial production growth and for inflation are improved by adjusting for a time-varying term premium. The adjusted term spread is given by

$$\hat{R}(n,t) - \hat{R}(m,t) = [R(n,t) - \hat{\Phi}(n,t)] - [R(m,t) - \hat{\Phi}(m,t)]$$

which can be rewritten using (4) as

$$\hat{R}(n,t) - \hat{R}(m,t) = \frac{1}{n}\sum_{i=0}^{n-1}\hat{R}(1,t+i) - \frac{1}{m}\sum_{i=0}^{m-1}\hat{R}(1,t+i) \qquad (5)$$

ie we are proposing the use of the term spread implied by the REHTS with expectations formed using a simple AR(*p*) model. The differences between the adjusted and observed term spread are shown in Graphs 3 and 4, where both of them are plotted for two- and three-year bonds as the long rate and the one-month interest rate as the short rate. Graph 3 presents the observed term spread together with the term spread implied by the adjustment with perfect foresight, ie replacing expected short rates with the actually realised one-month nominal rate.[8] The discrepancies between both measures are more extreme in the pre-EMU period, where the level and dynamics of the observed term spread are interpreted mainly as premium dynamics when using the adjustment method. The same qualitative conclusion applies if the three-month interest rate is used as the short rate. Graph 4, on the other hand, presents the observed term spread and the term spread implied by the adjustment using expectations formed by means of an AR(*p*) model. Due to the fact that the simple expectation-formation mechanism tended to overestimate the decrease of the nominal short-term interest rate in the pre-EMU period, the resulting synthetic long rates are very low compared to the one-month interest rate. This implies that a negative term spread prevails for the whole pre-EMU sample, which only turns positive at the end of 1999.

The potential improvement in the predictive content of the term spread for future developments in real activity and inflation will be tested and measured in the framework of an out-of-sample forecasting exercise for the growth rate of industrial production as well as headline and core inflation in the euro area.[9]

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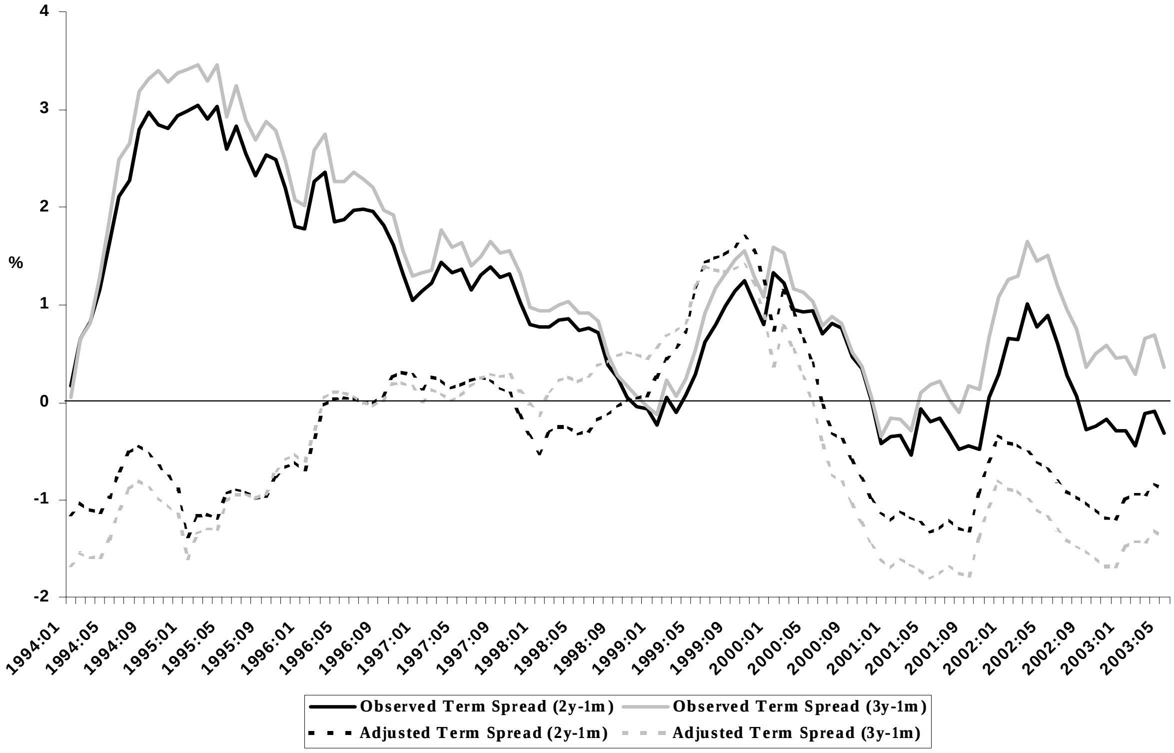
[8] The adjusted long rate for the last part of the sample was computed using simple projections of the short-term interest rate using all the available data.

[9] We will thus only consider what Estrella et al (2003) label a “continuous model”, as opposed to a “binary model”, with the latter aiming exclusively at forecasting the occurrence of recessions or the direction of change in inflation rates. Estrella et al (2003) provide evidence that binary models are more stable than those offering point forecasts of real activity. The choice of a continuous type of model for our exercise is conditioned by the fact that only one single recessionary episode has been observed in the aggregate euro area since 1990.

Graph 3

Observed and adjusted term spreads, perfect foresight

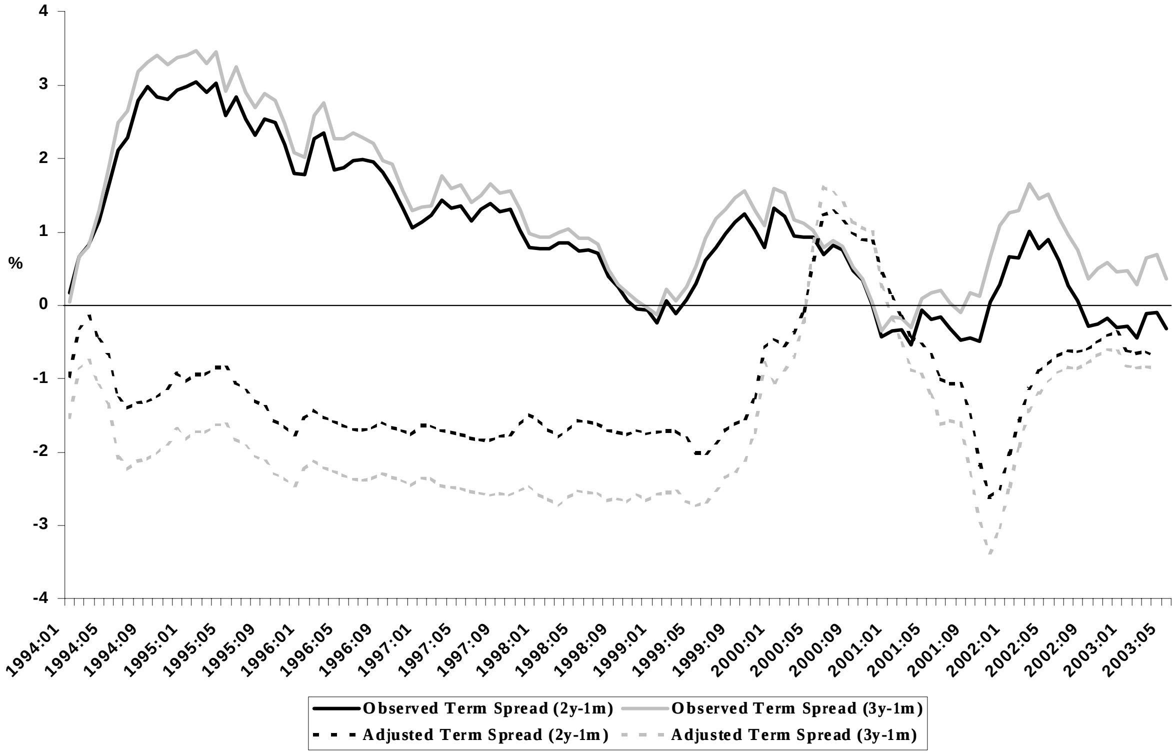
Long rate: two and three years; short rate: one month



Graph 4

Observed and adjusted term spreads, autoregressive expectations

Long rate: two and three years; short rate: one month



We will consider simple autoregressive distributed lags (ARDL(p,q)) models for forecasting industrial production growth and the inflation rate. For a given forecasting horizon h, the models estimated and used in the forecasting exercise are of the type

$$y_{t+h} = \delta + \sum_{i=0}^{p} \alpha_i y_{t-i} + \sum_{j=0}^{q} \beta_j x_{t-j} + \varepsilon_t \tag{6}$$

where y_t will alternatively be the yearly growth rate of industrial production or the inflation rate for the euro area. For a given dependent variable, x_t will alternatively be the observed and adjusted spread and ε_t is an iid random error with constant variance.

The forecasting exercise is carried out as follows. For a given value of the forecasting horizon, h, equation (6) is estimated using data up to period T using the observed spread as the x variable. With the estimated model, an h-steps-ahead out-of-sample forecast is generated. The observations for period $T+1$ are added to the estimation sample, (6) is re-estimated, and another h-steps-ahead forecast is computed. This is repeated until forecasts are obtained for all available observations of industrial production growth or the inflation rate since period $T+h$. The same procedure is then repeated for the adjusted spread as an x variable in (6). Notice that the adjustment procedure with AR(p) forecasts as expectations for the short rate which was described in the preceding section only requires data up to time t in order to obtain an estimate of $\Phi(n,t)$. The adjusted term spread assuming perfect foresight, however, uses future information for the adjustment method, so the results concerning this variable do not fulfil the usual requirements of a proper out-of-sample forecasting exercise, but are presented here for obvious comparison reasons.

The predictive ability of the different models used in the analysis will be compared in terms of root mean square forecasting error (RMSE). The h-steps-ahead RMSE of the model including variable x is given by

$$RMSE(x,h) = \sqrt{\frac{1}{N} \sum_{n=T+h}^{T+h+N} \left(y_n^{x,h} - y_n \right)^2}$$

where $y_n^{x,h}$ is the forecast of y_n obtained by the model with variable x and data ranging up to $T+n-h$, and N is the number of out-of-sample forecasts carried out. The Diebold-Mariano (Diebold and Mariano (1995), henceforth DM) test, which is described in the Appendix, will be used to compare the predictive accuracy of the models with the observed and adjusted term spread.

The results of the forecasting exercise for the rate of growth of industrial production are presented in Table 1. The procedure described above was carried out for adjusted and unadjusted term spreads with the two- and three-year bond as the long rate and the one- and three-month interest rate as the short rate. The lag lengths of the estimated ARDL(p,q) models are allowed to change with each new observation added to the in-sample period. In each replication, the lag lengths (p,q) chosen are the ones that jointly minimise the Schwarz criterion among those in the set $\{0,1,\dots,6\} \times \{0,1,\dots,6\}$. Table 1 reports the results of the forecasting exercise for forecasting horizons from six months to two years ahead, at six-month steps. In all cases, the first in-sample period was January 1994-January 1998, and forecasts were computed up to December 2002, the last observation of annual industrial production growth available. The last row of Table 1 presents the forecasting results for a simple autoregressive (AR) process, which is the natural benchmark of comparison if we want to evaluate the predictive content of the term spread in models such as (5).[10] The AR process is defined like in (6) without the second summation term on the right-hand side. The DM test statistic is provided in the table for those models that show better predictive abilities than the benchmark, and refers to the test for equal predictive accuracy against the AR model.

The results for the observed term spread confirm and expand the conclusions in Berk and van Bergeijk (2000, 2001). The simple AR model, which excludes the information contained in the term spread, performs better than the models including the unadjusted yield curve information in terms

[10] The procedure based on the Schwarz criterion was also used for choosing the optimal lag length for the AR process in each period. Qualitatively, the results remain unchanged if an unconstrained vector autoregression (VAR) using inflation and output growth data is used as the benchmark model. At most forecasting horizons, the simple AR model actually outperforms the VAR model in terms of forecasting error for output growth and inflation.

of RMSE for all forecasting horizons with the exception of two-years-ahead forecasts. For this forecasting horizon, only the model containing the term spread between the two-year bond and the three-month interest rate obtains a marginal improvement in the RMSE compared to the AR model, which is, however, insignificant according to the DM test.

Table 1

Forecasting comparison: industrial production growth

		RMSE			
		6 months	**12 months**	**18 months**	**24 months**
Adjusted spread (perfect foresight)					
Long rate	**Short rate**				
2 years	1 month	1.69 (–2.43***)	2.38 (–1.38*)	3.23	4.45
	3 months	1.71 (–2.09**)	2.29 (–1.61*)	3.36	4.03
3 years	1 month	1.73 (–2.03**)	2.08 (–1.87**)	2.90 (–0.01)	4.77
	3 months	1.72 (–2.07**)	1.99 (–2.06**)	2.95	4.56
Adjusted spread (AR expectations)					
2 years	1 month	2.67	2.63 (–0.46)	1.78 (–2.45***)	2.62 (–1.31*)
	3 months	2.65	2.67 (–0.32)	1.76 (–2.47***)	2.57 (–1.29*)
3 years	1 month	2.75	2.95	1.91 (–2.94***)	2.57 (–1.40*)
	3 months	2.73	2.97	1.89 (–2.82***)	2.63 (–1.52*)
Observed spread					
Long rate	**Short rate**				
2 years	1 month	2.82	3.53	4.05	6.07
	3 months	2.27	3.40	3.60	3.06 (–0.08)
3 years	1 month	2.78	3.50	4.35	5.83
	3 months	2.27	3.44	3.56	3.14
Benchmark AR model		2.09	2.82	2.91	3.07

Note: Numbers in parenthesis refer to the DM test statistic of the corresponding model against the AR model, asymptotically standard normal distributed. * (**) [***] refers to significance at 10% (5%) [1%] significance level.

While the results for the observed spread caution against the use of the information contained in the yield curve when forming predictions for real activity developments in the euro area, the forecasting exercise reaches a very different conclusion for the adjusted term spread. For forecasting horizons up to and including one year, the models including the premium-adjusted term spread with perfect foresight uniformly outperform all other models, independently of the interest rates used as long and short rates in the computation of the spread. The results of the DM test against the AR model conclude that the observed difference in predictive ability is significant in all cases. The predictive content of the adjusted term spread with perfect foresight ceases to exist, however, for longer forecasting horizons. For 18-months-ahead predictions, only one of the models with adjusted term spreads and perfect foresight presents an insignificantly lower forecasting error than the AR model, and for the two-year forecasting horizon, all models including the adjusted term spread are outperformed by the minimal benchmark AR model.

The improvement in the predictive ability of the premium-adjusted term spread with perfect foresight is not surprising, as it includes actual information on the development of short-term interest rates in the out-of-sample period. The forecasts obtained from the premium-adjusted term spread using AR(p) expectations, by contrast, are based exclusively on in-sample data. The results for long-term forecasts with the model containing the adjusted term spread using AR(*p*) expectations indicate an overwhelming improvement of the prediction error for forecasting horizons higher than a year ahead. Independently of the rates used to form the term spread, all models including this variable outperform significantly the benchmark model at 18- and 24-months-ahead horizons, with reductions of the RMSE up to 40% compared with the simple AR model and 55% if compared to the model including the observed spread. The fact that the forecasting horizon where improvements are significant has shifted forward as compared to the perfect foresight case is explained by the relatively high inertia of the autoregressive forecasts (changes in direction of the trend which is estimated when forming expectations tend to be picked up with around 12 months' delay).

The results are very different if the variable to be predicted is inflation. Table 2 presents the results for the headline inflation rate in the euro area (defined as yearly change in the harmonised index of consumer prices), and Table 3 presents the results for the core inflation rate (defined as yearly change in the harmonised index of consumer prices excluding energy and unprocessed food).

Table 2

Forecasting comparison: headline inflation

		RMSE			
		6 months	**12 months**	**18 months**	**24 months**
Adjusted spread (perfect foresight)					
Long rate	**Short rate**				
2 years	1 month	0.58	1.07	1.76	2.29
	3 months	0.59	1.06	1.68	2.26
3 years	1 month	0.55	1.00	1.53	2.22
	3 months	0.55	0.99	1.55	2.21
Adjusted spread (AR expectations)					
2 years	1 month	0.48	0.89 (–0.71)	1.64	2.44
	3 months	0.46 (0.52)	0.76 (–1.33*)	1.40	2.25
3 years	1 month	0.49	2.33	1.67	2.39
	3 months	0.48	0.91 (–0.24)	1.52	2.27
Observed spread					
Long rate	**Short rate**				
2 years	1 month	0.55	0.98	1.43	2.16
	3 months	0.51	1.15	1.59	2.12
3 years	1 month	0.56	1.02	1.46	2.18
	3 months	0.51	1.10	1.57	2.15
Benchmark AR model		0.48	0.96	1.24	1.56

Note: Numbers in parenthesis refer to the DM test statistic of the corresponding model against the AR model, asymptotically standard normal distributed. * refers to significance at 10% significance level.

Although the adjusted term spread using AR(*p*) expectations achieves lower forecast errors than all other models in some cases for forecasting horizons up to one year, only the model with the adjusted two-year–three-month spread is able to outperform the benchmark significantly for one-year-ahead predictions. Neither the information contained in the observed term spread nor that contained in the adjusted term spread with perfect foresight improves the predictions on inflation based on its own past history at any forecasting horizon.[11]

Table 3

Forecasting comparison: core inflation

		RMSE			
		6 months	**12 months**	**18 months**	**24 months**
Adjusted spread (perfect foresight)					
Long rate	**Short rate**				
2 years	1 month	0.35	0.77	1.17	1.46
	3 months	0.35	0.75	1.17	1.46
3 years	1 month	0.28 (–0.67)	0.45 (–0.97)	0.89 (–0.88)	1.23 (–1.39*)
	3 months	0.28 (–0.58)	0.48 (–0.80)	0.93 (–0.77)	1.26 (–1.21)
Adjusted spread (AR expectations)					
2 years	1 month	0.52	1.29	0.93 (–1.61*)	1.00 (–1.24)
	3 months	0.49	1.36	0.92 (–1.16)	0.89 (–2.04**)
3 years	1 month	0.58	2.18	1.78	0.99 (–1.16)
	3 months	0.54	1.66	1.89	0.92 (–1.72**)
Observed spread					
Long rate	**Short rate**				
2 years	1 month	0.46	0.95	1.03 (–0.96)	1.21 (–2.19**)
	3 months	0.41	0.82	1.20	1.66
3 years	1 month	0.45	0.89	1.00 (–1.34*)	1.21 (–2.35**)
	3 months	0.41	0.78	1.27	1.64
Benchmark AR model		0.34	0.68	1.13	1.39

Note: Numbers in parenthesis refer to the DM test statistic of the corresponding model against the AR model, asymptotically standard normal distributed. * (**) refers to significance at 10% (5%) significance level.

However, the term spread, in both its adjusted and unadjusted form, seems to be partly useful for obtaining forecasts of core inflation. The results in Table 3 show that the models including the observed term spread with the one-month interest rate significantly outperform the benchmark model in predicting core inflation rates at long horizons. The improvement is still greater if the adjusted spread with AR(p) expectations is used, with reductions of the RMSE over the benchmark of more than 35%. The model with the adjusted term spread using the difference between the adjusted two-year bond rate and the adjusted three-month interest rate presents the best forecasting abilities at the two-years-ahead horizon, and outperforms (with a DM test statistic of ˜ 1.71) the best model among those using the observed spread. Surprisingly, marginal improvements over the benchmark are observed for the adjusted term spread with perfect foresight only for two-years-ahead forecasts, and these are of a small magnitude compared to the improvements obtained using the adjustment with AR(p) expectations.

[11] Estrella et al (2003) note that the relationship between real activity and the term spread is of a more stable nature than that between inflation and the term spread. Our results for the inflation rate may as well reflect the existence of one or more structural breaks in the underlying data-generating process.

Given the way in which the adjustment takes place with AR(*p*) expectations, the adjusted term spread is computed using exclusively information on the short-term interest rate. The results presented above could thus be interpreted as evidence that the predictive power of the term spread is determined by the dynamics in the short-term rate. The aggregation implied by the REHTS is, according to the results presented, a useful way of disentangling the part of the term spread whose dynamics actually contain information on future macroeconomic developments. If the adjustment method is to be relied upon, one would expect that no significant information on future developments in real activity and inflation should be present in the risk premia estimates plotted in Graph 2. Table 4 presents the results of the forecasting exercise explained above using the risk premia implied by the decomposition with AR(*p*) expectations as the *x* variable.

Table 4

Forecasting comparison results for risk premia estimates

Risk premia		**RMSE**			
		6 months	**12 months**	**18 months**	**24 months**
Industrial production growth					
Long rate	**Short rate**				
2 years	1 month	2.78	4.37	4.35	4.21
	3 months	2.80	4.52	3.83	3.73
3 years	1 month	2.81	4.74	4.54	4.14
	3 months	2.81	5.22	3.99	3.65
Headline inflation					
2 years	1 month	0.50	0.94 (0.14)	2.45	2.34
	3 months	0.70	1.08	1.59	2.16
3 years	1 month	0.50	0.94 (0.17)	1.73	2.25
	3 months	0.67	1.10	1.59	2.38
Core inflation					
2 years	1 month	0.45	0.88	1.26	1.81
	3 months	0.42	0.78	1.47	1.32 (0.95)
3 years	1 month	0.45	1.07	1.34	1.50
	3 months	0.44	0.81	1.55	1.41

Note: Numbers in parenthesis refer to the DM test statistic of the corresponding model against the AR model, asymptotically standard normal distributed.

The results in Table 4 present the RMSE obtained in the forecasts when using the risk premium with respect to the one- and three-month interest rate as explanatory variables in the out-of-sample exercise presented above. There is no improvement on the models where industrial production growth, headline inflation or core inflation are explained by their own past for any forecasting horizon and any risk premium estimate. These results indicate that the decomposition used tends to be successful in isolating the part of the term spread with predictive properties for industrial production growth and, notwithstanding the limitations of this link, also with inflation.

The method used to adjust the term spread for time-varying risk premia renders an adjusted term spread composed exclusively of autoregressive expectations on the short rate, which are aggregated according to the REHTS using (5). Whether imposing the structure implied by (5) actually improves the forecasting abilities of the term spread as compared to using exclusively the information embodied in the short rate data without the restrictions implied by the aggregation method can also be checked empirically. Table 5 presents the results of the forecasting exercise using the monthly change in the short rate as an explanatory variable in (6).[12] There is no evidence of significant improvement over the forecasts of the benchmark model for any variable at any forecasting horizon. The results for the short rate can be interpreted as a robustness check of the simple methodology proposed, and they draw attention to the empirical relevance of the method of aggregation of expectations implied by the REHTS when assessing the predictive abilities of the term spread for output growth and inflation.

Table 5

Forecasting comparison results for the short rate

RMSE			
6 months	**12 months**	**18 months**	**24 months**
Industrial production growth			
2.11	2.84	2.86 (–0.54)	3.23
Headline inflation			
0.50	0.87 (–0.89)	1.57	2.27
Core inflation			
0.36	0.73	1.15	1.36 (–1.04)

Note: Numbers in parenthesis refer to the DM test statistic of the corresponding model against the AR model, asymptotically standard normal distributed.

5. Conclusions and paths of further research

This paper reinvestigates the informational content of the yield spread for real activity and inflation for the euro area aggregate. The motivation is threefold. First, at the theoretical level, a number of possible channels have been put forward in the literature that would suggest a systematic empirical relationship between the yield spread and current and/or future real activity. Second, at the level of data availability, four and a half years of genuine euro area data make it worthwhile to investigate the issue empirically. Third, previous research has not paid attention to the substantial difference of the monetary policy regime in place prior to the start of EMU, which may have strongly influenced risk premia over time. Contrary to previous research on the euro area, this paper explicitly pays attention to disturbances of the term spread from time-varying risk premia. We put forward a simple, purely empirical adjustment procedure for a time-varying term premium based on the rational expectations hypothesis of the term structure, and find that significant improvements can be achieved in the predictive content of the term spread if the dynamics of the risk premium are taken into account in its computation.

The results of a forecasting exercise using adjusted and unadjusted term spreads show that, for the euro area aggregate, modelling the risk premium adequately is a necessary requirement in order to exploit the information embodied in the term spread for predictions in the development of real activity and inflation. Regarding real activity, of all possible models including the term spread, only those where the adjustment was performed were able to deliver significantly better medium-run forecasts than simple models where the growth rate of industrial production is explained by its own past history.

[12] Augmented Dickey-Fuller tests could not reject the existence of a unit root in the series of one-month rates at any reasonable significance level.

For forecasting horizons exceeding one year, the models including the premium-adjusted term spread, where the expectations on the short rate are modelled through a simple autoregressive model, uniformly outperform all other models. This result arises independently of the interest rates used as long and short rates in the computation of the spread. For the case of inflation, however, the results are more mixed, but evidence of improvement in the forecasting abilities of the term spread after the premium adjustment was provided for two-years-ahead forecasts of core inflation.

We conclude that, if distortions arising from time-varying risk premia are filtered out, the term spread can - despite the substantial limitations imposed on econometric estimates by the necessity to use *synthetic* pre-EMU data - nevertheless serve as one useful indicator (among others) to gauge future developments in real activity and, to a lesser extent, (core) inflation. In this sense, it seems worth monitoring as part of the "economic analysis" within the framework of the Eurosystem's monetary policy strategy. In particular, after adjusting for the existence of a time-varying risk premium, the term spread could be useful in order to check the robustness of forecasts produced by more extensive macroeconomic models.

An alternative reading of our results is that - for the euro area - using information embodied in short-term interest rates yields better forecasting results for both real activity and (core) inflation than the term spread. In other words, the medium-term end of the yield curve used in our study seems to contain no additional information. However, our results show that the aggregation of expectations on short rates implied by the REHTS seems to play an important role in the predictive properties of the adjusted term spread. This interpretation would raise serious questions about the widespread reference by financial analysts and policy commentators to the (term-spread-unadjusted) yield curve as a market expectations indicator.

Finally, it may also be that the policy regime break induced by the inception of EMU pollutes empirical analysis at this stage too much. In this case, the issue might be resolved over time, as longer time series become available and the regime break becomes an event which is only relevant for the beginning of the sample. Linked to that, it may also be that the use of more sophisticated econometric methods will in the future be able to shed some light on the reasons for the predictive failure of the observed spread in the euro area.

In this vein, Venetis et al (2003) provide evidence concerning the existence of threshold effects in the relationship between the term spread and real activity for Canada, the United Kingdom and the United States. The use of non-linear time series models to assess the informational content of the term spread on future developments in real activity can thus be seen as a possible avenue of future research in order to provide further evidence on the leading indicator properties of the slope of the yield curve.

Appendix

Data sources

- One-month interest rate, euro area aggregate. Source: Datastream.
  Range: November 1990-May 2003.
- Three-month interest rate, euro area aggregate. Source: Bank for International Settlements.
  Range: January 1990-May 2003.
- Two-year bond yield, secondary market, benchmark, euro aggregate. Source: Bank for International Settlements.
  Range: January 1994-May 2003.
- Three-year bond yield, secondary market, benchmark, euro aggregate. Source: Bank for International Settlements.
  Range: January 1994-May 2003.
- Industrial production index, euro aggregate, seasonally adjusted. Source: Eurostat.
  Range: January 1990-December 2002.
- Harmonised index of consumer prices (all items and all items excluding energy and unprocessed food), euro aggregate, seasonally adjusted using the Census X12 method. Source: European Central Bank.
  Range: January 1990-May 2003.

The Diebold-Mariano test for comparing predictive accuracy

The DM test is an asymptotic test for the null of equal predictive accuracy of two models. In the framework proposed above, consider two models using variables x_1 and x_2 respectively. For a given forecasting horizon h, the null hypothesis in the DM test is that

$$d_n = E\left[g(e_{1n}) - g(e_{2n})\right] = 0$$

where e_{1n} is the forecasting error produced by the model with variable x_1 when forecasting Δy_n (that is, $e_{1n} = \Delta y_n^{x_1,h} - \Delta y_n$), e_{2n} is defined analogously for x_2, and $g(z)$ is a prespecified loss function associated with the forecast error. In our case, the loss function is a quadratic one, so that $g(z) = z^2$. The DM test is based on the observed average forecast error difference, $\bar{d}$, The DM test statistic is given by

$$S_1 = [\hat{V}(\bar{d})]^{-1/2}\bar{d}$$

where $\hat{V}(\bar{d})$ is an estimate of the asymptotic variance of $\bar{d}$, given by

$$\hat{V}(\bar{d}) = \frac{1}{N}\left(\hat{\gamma}_0 + 2\sum_{k=1}^{h-1}\hat{\gamma}_k\right)$$

where $\hat{\gamma}_k$ is the k-th order sample autocovariance of the forecasting error difference series. The asymptotic distribution of S_1 is standard normal, so tests for equality of predictive accuracy between different models can be easily carried out.[13]

[13] The DM test methodology is not free of criticism. For a recent critical assessment of testing predictive accuracy using the DM test statistic, see Kunst (2003).

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