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FORECASTING OUTPUT GROWTH AND INFLATION IN THE EURO AREA: ARE FINANCIAL SPREADS USEFUL?

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Abstract

This paper deals with the usefulness of several measures of financial spreads (the slope of the yield curve, the reverse yield gap, the credit quality spread) for forecasting real economic activity and inflation in the euro area. A quarterly Bayesian vector autoregression model is used to assess the marginal forecasting power of financial spreads for real economic activity and inflation. A benchmark BVAR is set up, containing real GDP, inflation and key indicators of monetary policy and foreign macroeconomic variables. The properties of the spreads as leading indicator are then assessed by augmenting the benchmark BVAR with the spreads, one at a time. We find that financial spreads have no or negligible marginal predictive content for either target variable. Overall, there is no ready-to-use financial indicator that can replace an encompassing multivariate model for the prediction of target variables in the euro area.

JEL classification: C11; C32; C53

Keywords: financial spreads, bayesian VAR models, bayesian analysis, forecasting

Contents

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1. Introduction[1]

There is a vast literature on the prediction of real output growth and inflation using financial spreads. The basic idea is that asset prices are formed by rational forward-looking agents in financial markets and may embody useful information on future real economic activity. In addition, spreads are readily observable, well in advance of macroeconomic data for the current period, and are not subject to revision. The indicator properties of financial spreads have been investigated in seminal works for the US (Laurent, 1988, 1989; Harvey, 1988, 1989; Stock and Watson, 1989; Bernanke, 1990; Mishkin, 1990, 1991; Estrella and Hardouvelis, 1991; Jorion and Mishkin, 1991; Friedman and Kuttner, 1991), and this approach has been extended to the UK and to other OECD countries (Davis, 1993; Davis and Henry, 1994; Plosser and Rouwenhorst, 1994; Davis, Henry and Pesaran, 1994; Bonser-Neal and Morley, 1997; Kozicki, 1997; Gerlach, 1997; Davis and Fagan, 1997; Estrella and Mishkin, 1997), including European ones.

The contributions that have used a bivariate approach to assess whether there is a significant correlation between spreads and future inflation or output growth have generally supported the idea that spreads may play an important role in improving forecasts. In particular, the spreads are often able to add information to a simple bivariate autoregression for output growth and inflation (Harvey, 1988; Estrella and Hardouvelis, 1991; Davis and Henry, 1994; Estrella and Mishkin, 1997). However, these relationships are not stable over time and lead to relatively poor out-of-sample forecasting performance.[2] In addition, the conclusions reached using the bivariate approach may no longer hold when additional variables are considered. In particular, in multivariate models the inclusion

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[2] For example, Haubrich and Dombrosky (1996) and Dotsey (1998) showed that the predictive content of the term spread for the US economic activity has diminished since 1985; Smets and Tsatsaronis (1997) concluded that the predictive content of the term spread is not time-invariant for output growth in Germany or the United States, while Davis and Fagan (1997), focusing on European countries, found that spreads only very rarely satisfy the three conditions of significance, stability and improved out-of-sample forecasts at once.

of some monetary policy variables, such as short-term interest rates and some monetary aggregates, as well as of other leading indicators may crowd out the spreads.[3]

To date the literature on the euro area as a whole has only focused on the forecasting properties of the slope of the yield curve. Berg and Van Bergeijk (2000) and Nicoletti Altimari (2000) found that no additional information is present in this spread beyond that contained in past values of inflation and output. At the same time, in the context of dynamic factor analysis, Altissimo et al. (2001) found that the slope of the yield curve is procyclical and leading for the euro area GDP, but the association with the business cycle is weaker than that of other leading variables. Cristadoro et al. (2001) found that this spread is lagging and negatively related to a core inflation index for the euro area.

This paper integrates previous ones along four dimensions. First, using aggregate euro area variables, it focuses on the marginal predictive content of several financial spreads, in addition to the slope of the yield curve: two versions of the reverse yield gap, measured by the difference between the bond yield and the yields on domestic equities (respectively, based on earnings and dividends) and a credit spread variable (defined as the difference between the bank retail interest rate on short-term loans to firms and the short-term interest rate).

Second, the empirical analysis is conducted in a multivariate framework, which enables us not only to investigate the marginal predictive content of financial spreads over and above the information contained in other macroeconomic variables, but also to overcome the critiques related to the choice of a single-equation approach. As is well known, this method treats all leading variables as predetermined, implying arbitrary restrictions on dynamic relationships between spreads and goal variables and ignoring more complex interactions, which may be captured using spreads in combination with other leading indicators. This choice is supported by the recent empirical results of Stock and Watson (2001) for the G-7 and Forni et al. (2002) for the euro area.

Third, techniques based on classical inference may be subject to limits in this context,

[3] See Kozicki (1997) and Estrella and Mishkin (1997) for some examples.

given the instability of the predictive equations involving spreads, and considering the small number of observations for the euro area time series. Thus, we adopt a Bayesian VAR framework, which provides a way to specify time-varying coefficient models and overcomes the over-parameterisation problem typical of standard VAR models, improving out-of-sample forecasting performance.

Fourth, most of the studies cited draw their conclusions entirely from in-sample statistics, while we propose results strictly based on out-of-sample forecasts. In addition, as the marginal predictive content of financial spreads may depend crucially on the information set used, all BVAR models are estimated recursively and the out-of-sample forecasts computed in the most efficient way using the information actually available to the forecaster, through a conditional forecasting approach. This is an attempt to avoid biasing the results in favour of indicators that tend to become available only with a substantial delay. For the euro area this is a relevant issue, as real GDP data are released with a lag of at least one quarter.

The paper is organized as follows. Section 2 focuses on the theoretical foundations of financial spreads. Section 3 describes the data. Section 4 points out the potential usefulness of financial spreads for both output growth and inflation by means of a cross-correlation analysis. Section 5 describes the BVAR methodology, the conditional forecasting approach and presents the empirical evidence and the main results of the out-of-sample forecasting exercise. Section 6 concludes.

2. Theoretical Foundation of Financial Spreads

Financial spreads are defined as differences in asset yields. The theory of finance suggests that all assets in the economy are imperfect substitutes for one other, mainly because of differences in liquidity, maturity and risk and covariances of yields on alternative assets. Changes in degree of substitutability can arise from structural factors, such as taxes and portfolio regulation, or from cyclical ones, such as monetary policy shifts. When the latter occur, the subsequent movements in financial spreads can provide information about future output growth and inflation. In this section we discuss the theoretical foundations of three different kinds of spreads, those most commonly used in the literature, namely the slope of the yield curve, the reverse yield gap and the banking spread.

The *slope of the yield curve* is the difference between a long-term and a short term interest rate. Its theoretical link with future output growth and inflation is given by the expectation hypothesis and the Fisher equation

$$i_t^m = \frac{1}{m}\sum_{k=0}^{m-1} E_t\left(i_{t+k}\right) + tp_t^{(k)} \tag{1}$$

$$i_t^m = rr_t^m + E_t\left(\pi_t^m\right) \tag{2}$$

Equation (1) states that the yield of a long-term risk-free asset of maturity m can be decomposed into the sum of the current yield and the weighted average of the expected future yields of a short-term risk-free asset, plus a term premium depending on the length of immobilization. As the short-term interest rates basically reflect the stance of monetary policy, a declining yield curve can be considered as signalling *currently* tighter monetary policy and, consequently, indicating a future slowdown in real economic activity. Secondly, this spread can be an indicator of the *expected* monetary policy. If market participants expect an expansive monetary policy, they believe that the future reduction in short-term rates will boost growth. At the same time, as they also expect higher inflation, higher than the expected reduction in the real interest rate, the long-term rate might increase and so would the spread. Through this mechanism, greater future growth would be correlated with a current increase in the spread. Equation (2) says that spot nominal interest rates contain at least two components, the m-period real interest rate, rr_t, and the expected rate of inflation m periods ahead. Therefore, under the restrictive assumptions of constancy of the real interest rate over time, absence of a term premium implied by the opportunity cost of immobilisation, and the validity of the expectations theory of the term structure, the slope of the yield curve can also provide a measure of the expected future path of inflation. All these explanations are derived from the role of monetary policy in affecting real economic activity via interest rates. Nevertheless, some other theoretical arguments have pointed out the existence of a positive relationship between the slope of the yield curve and future economic activity. Harvey (1988) developed a CCAPM model that yields this relationship based on the smoothing of consumption implied by the traditional Euler equation. Kydland and Prescott (1988) set up an RBC model where again consumption smoothing leads to the same first order condition.

The *reverse yield gap* is defined as the difference between yields on long-term or short-term government securities and the dividend yield on domestic equity. Securities issued by the government are essentially risk-free assets and also have a fixed coupon and expiration date, whereas equities are risky assets and give only variable dividend yields; accordingly, this spread reflects the premium that an investor is likely to demand to compensate for the extra risk.[4] A market index of yields, as opposed to yields on individual bonds, is essentially a weighted average of all the expectations of market participants on the future default risk of the economy, when all the idiosyncratic risk components associated with the balance sheets of each individual company or bank have been valued. Thus, to the extent that such expectations are accurate, increases in this spread will predict downturns in economic activity. In addition, the relation of the reverse yield gap to inflation might be expected to be positive, because a rising spread will accompany a tightening of monetary policy in response to increased inflationary pressures, which in turn also raises the default risk.

Finally, the *bank spread* is defined as the difference between a banking system lending rate and the short-term interest rate that reflects the monetary policy stance. The leading properties of this spread basically reflect the credit channel in the transmission of the monetary policy and are the more relevant, as more the investment and consumption decisions of non-financial agents depend on the banking system. The economic literature identifies two different channels of transmission, the *bank lending channel* and the *balance sheet channel*. The first is linked to the fact that, after a contractionary monetary policy, the subsequent reduction in the amount of deposits leads to changes in the banking system's outstanding liabilities, which has to be counterbalanced by a recomposition of the assets outstanding by means of a contraction of holdings of securities and loans to customers. With a low degree of access to capital markets, the fall in the supply of

[4] In addition, the differential between the yield on a private bond and a public bond of the same maturity, callability and tax features can be also seen as an indicator of the market's assessment of default risk.

credit is greater than that of the securities. This channel is stronger in the presence of forms of credit rationing (Stiglitz and Weiss, 1981; Blinder, 1987). The balance sheet channel can be explained by means of the effects of the changes in relative prices of financial assets held by the economic agents demanding the loans. After a monetary tightening the fall in the price of debt securities and equities leads to a reduction in the value of agents' collateral (Gertler and Hubbard, 1988), which is greater if the firms with no access to the equity market are considered riskier; therefore, in the presence of asymmetric information between lenders and borrowers, there will be increased adverse selection and moral hazard in bank lending towards firms with reduced net worth. The banking system becomes more *flight to quality* oriented (Bernanke, Gertler and Gilchrist, 1994; Lang and Nakamura, 1995) and firms with greater difficulty in obtaining credit decide to postpone or scale down investment plans. In addition, the banking spread may depend on imperfect substitutability between financial instruments. If agents expect a recession and fear being short of liquidity, they will borrow more, leading to changes in the macroeconomic financial structure (Kashiap, Stein and Wilcox, 1992). If the gap between the supply and demand of financial instruments varies differently across markets over the business cycle, then variations in banking spreads may carry information about the future business cycle.

3. The Data

We use quarterly time series from 1980q1 to 2002q4; all data are seasonally adjusted (except interest rates and financial spreads), and in natural logs (except these two and the euro-dollar exchange rate). The set of euro-area variables considered are real GDP, HCPI inflation, various measures of the money stock (M1, M2 and M3), the short-term (three-month) interest rate, two measures of international competitiveness - the real effective exchange rate (CPI-based) and the nominal exchange rate of the euro with the dollar. In addition, four financial spreads are considered: first is the slope of the yield curve, defined as the difference between the 10-year government bond yield and the 3-month money market interest rate; second is the reverse yield gap, defined as the difference between the long-term interest rate and yields on domestic equity market. We consider two versions of this indicator, using the dividend yields and the earning yields on the euro area equity market. These two versions could have different leading properties, as earnings reflect actual cash flow whereas dividends, which result from corporate decisions, may be an important indicator of mangers' views, also reflecting their private information. Third is the banking spread, defined as the difference between bank retail interest rate on short-term loans to firms and the short-term interest rate. In this paper, we essentially focuses on a *credit quality spread*, which basically expresses the banks' own opinion about the default risk and, consequently, on the future state of the business cycle.[5] For the pre-EMU period (1980q1 to 1998q4) all euro-area variables are obtained via aggregation of their national-level counterparts. Fourth, some foreign indicators are also considered, in order to control for demand and supply side pressures from international markets: the commodity price index, the oil price, real US GDP and short-term (three-month) US interest rates. For details on sources oand the construction of the variables, see the Appendix I. In Figure 1 we report the historical pattern of four financial spreads and the target variables, namely, four-quarter growth rates of real GDP and four-quarter inflation rates.

4. Cross-Correlation Analysis

In this section we run a preliminary investigation on the information content of each financial spread for both real output growth and inflation. To this end, we compute the cross-correlation coefficients among variables in order to establish both pro-cyclical or counter-cyclical and leading or lagging properties. We define a financial variable as leading (lagging) if the maximum cross-correlation coefficient, in absolute value, corresponds to a lag (lead) of the variable relative to the contemporaneous value of t. The series is synchronous if the maximum absolute value is the contemporaneous correlation. Finally, the

[5] In theory it is possible to construct other bank spreads. The *banking yield curve spread* is defined as the difference between two bank rates of different maturities (for example the long-term credit rate minus the prime rate or the mortgage rate). The *intermediation margin spread* is defined as the difference between the prime rate or average rate and the call money rate, which is considered here as the marginal cost of finance for banks. An increase in the intermediation margin may stimulate the supply of credit by banks and therefore economic activity, especially when some categories of agents, such as households and small firms, are credit-rationed and do not have access to alternative means of external finance. Nevertheless, because of the lack of sufficiently long time series for the euro area, we could not consider these alternative spreads in the empirical analysis.

series is classified as procyclical if the contemporaneous correlation coefficient is positive and counter-cyclical if it is negative. Real GDP time series is expressed in four-quarter growth rates, while the inflation rate is computed as the four-quarter change in the level of the price index. Interest rates and financial spreads are expressed in levels. As is well known, growth rates overemphasize high frequencies and under-weight low ones, so, we opted for showing the results obtained at the business cycle frequencies using filtered time series with a band-pass filter proposed by Baxter-King (1999). Following the suggestion of Agresti and Mojon (2001), for all the euro-area variables we allow the upper bound on the length of the business cycle to be 40 quarters (10 years) instead of 32 (8 years). In addition, the truncation of the band pass filter is done with 8 leads and lags (instead of 12), as the series we considered starts in the 1980's.[6]

We also focused on two relevant sub-sample periods, namely 1994q1-2002q4 and 1999q1-2002q4. The quarter 1994q1 represents the beginning of Stage Two of Economic and Monetary Union (EMU) whith the establishment of the European Monetary Institute (EMI). This phase strengthened central bank cooperation and monetary policy coordination and made the preparations for the establishment of the Eurosystem. The quarter 1999q1 marked the beginning of Stage Three of EMU and was characterized by the transfer of the power to set the single monetary policy to the ECB and the introduction of the euro as single currency. We expect that gradual and sweeping changes in monetary policy conditions can be reflected in the leading properties of the spreads for euro area as a whole.

Looking at the data (see Table 1), we see that the business cycle has been characterized by a strong reduction in the volatility of inflation between 1980q1 and 2002q4. The standard deviation of the harmonized consumer price index (HCPI) fell to less than half that of real GDP. The short-term and the long-term interest rates are in general more volatile than real GDP, although in the last few years the long-term interest rates seems to be smoother. All financial spreads except the credit quality spread are more volatile

[6] In order to provide further robustness checks that the key result we found does not depend on the filter used, we also repeated the same analysis using the traditional Hodrick-Prescott filter, with constant weighting of 1600, and another Baxter and King band pass filter, where the "longest" business cycle period is limited to 8 years (as in Stock and Watson, 1999). Differences in the results are negligible.

than real GDP. Figure 2 compares business cycle components of financial spreads and those of real GDP over the entire sample period.

In Table 2 and 3 we report the cross correlations between financial spreads and both real GDP and inflation. The results are consistent with the economic theory and show some important differences in the indicator properties of the various spreads.

Over the entire sample period, both short-term and long-term interest rates are procyclical and lead GDP slowdowns by of four to eight quarters. Short-term rates show the highest correlation with real GDP at a horizon of one year, long-term, at six quarters. The slope of the yield curve is counter-cyclical and leads an economic recession by two to four quarters. Both measures of the reverse yield gap are strongly procyclical and lead real GDP slowdowns by six or eight quarters. The credit quality spread is strongly counter-cyclical and leads to real GDP slowdown by about six quarters. This pattern also holds over the most recent sub-sample periods, but the indicator properties of financial spreads for real economic activity have improved remarkably over time. Of the four financial spreads, the slope of the yield curve appears to have the best indicator properties with future real GDP at a horizon of one year, while the short-term interest rate still exhibits a higher correlation with future economic activity at horizons from four to six quarters ahead. This feature is probably consistent with the view that the level of the term structure is more informative than its slope. At longer horizons the reverse yield gap and the credit quality spread are characterized by indicator properties comparable to those of the long-term interest rate.

With respect to inflation, the analysis conducted over the entire sample period shows that, at shorter horizons (up to two quarters ahead) the short-term interest rate is the best leading indicator for inflation. At somewhat longer horizons, four to eight quarters ahead, the long-term rate substantially outperforms all other financial variables. Among financial spreads, only the reverse yield gap shows good leading indicator properties. Nevertheless, in most recent times both the slope of the yield curve and the two measures of the reverse yield gap seem to increase their leading properties. Notice that over the period 1994q1-2002q4 the correlation between the slope of the yield curve and inflation is highest between six and eight quarters ahead, but over the period 1999q1-2002q4 it declines significantly. By contrast, the credit quality spread is completely useless in forecasting inflation at all horizons and over the entire period.

In summary, our univariate analysis supports the idea that financial spreads are useful leading indicators of both real GDP and inflation in the euro area, even if their predictive power is generally not stable over time.

5. Multivariate Analysis

5.1. The BVAR Approach

To investigate the usefulness of financial spreads in forecasting output growth and inflation, we model the euro area economy in a Bayesian VAR framework. Let us consider the state space representation of an unrestricted VAR model of order p

$$y_t = A_t(L)y_{t-1} + C_t w_t + \epsilon_t \tag{3}$$

$$\epsilon_t \sim N(0, \Sigma) \tag{4}$$

$$\beta_t = G\beta_{t-1} + F\overline{\beta} + \eta_t \tag{5}$$

$$\eta_t \sim N(0, \Omega_t) \tag{6}$$

where the dependent variable y_t is an $(n \times 1)$ vector in a VAR, w_t is an $(r \times 1)$ vector of deterministic variables (constant term, trends, dummies), $\beta_t = vec\left[A_t(L), C_t(L)\right]$ is the stacked version of the coefficient vector, G and F are $(nk \times nk)$ matrices and $k = np + r$ is the number of coefficients in each equation of the VAR model. Equation (5) is the most general law of motion of the coefficient vector in a time-varying framework, which can allow for both stationary and non-stationary environment.[7] Time variation is allowed in order to consider the presence of relevant breaks in the quarterly time series. As the number of parameters increases rapidly with the number of variables and the lag order, the specification of VAR models is constrained by a "degrees of freedom" problem, leading

[7] For example, if the roots of G are stationary the law of motion of the coefficients displays reversion toward the mean $\overline{\beta}$.

to poor forecasting performance due to over-parameterisation, even though the within-sample fit may be good. The Bayesian VAR approach, developed by the seminal papers of Doan, Litterman and Sims (1984) and Litterman (1986), reduces the dimension of the parameter space by specifying a prior distribution for the parameters. In this framework the law of motion of the coefficient is treated as the first layer of a hierarchical prior, while the specification of the mean and the variances of the distribution of the coefficients are basically the features of the second layer. Specifically, we consider the following simplified prior structure, specified in terms of a small set of hyperparameters

$$G = \lambda_0 \cdot I \qquad F = I - G \tag{7}$$

$$\overline{\beta}_{ij,l} = \left\{ \begin{array}{ll} 1 & i = j, \quad l = 1 \\ 0 & otherwise \end{array} \right\} \tag{8}$$

$$\Omega_t = \lambda_1 \cdot \Omega_0 \tag{9}$$

$$\Omega_{0ij,l} = \left\{ \begin{array}{ll} \lambda_2 \cdot l^{-\lambda_3} & i = j, \quad \forall l, \quad i,j = 1,...,n \\ \lambda_2 \cdot f(i,j) \cdot l^{-\lambda_3} \cdot (\widehat{\sigma}_i / \widehat{\sigma}_j)^2 & i \neq j, \quad \forall l, \quad i,j = 1,...,n \end{array} \right\} \tag{10}$$

$$\Omega_{0ik} = \lambda_2 \cdot \lambda_5 \cdot (\widehat{\sigma}_i)^2 \qquad k = 1,...,r \tag{11}$$

According to (7) and (9), the law of motion for the parameters vector is specified as a first-order autoregressive process with decay toward the mean, in which the hyperparameter λ_0 controls the extent of the mean reversion of the coefficient vector, while the hyperparameter λ_1 measures the degree of time variation in the law of motion of the coefficients. In particular, when simultaneously $\lambda_0 = 1$ and $\lambda_1 = 0$ there is no time variation in the model.

According to (8), in each equation the coefficient vector is assumed to be an independent and normal prior distribution with zero mean, except for the coefficient on the first lag of the own variable which has a mean of unity (the so called "Minnesota prior"), based on the assumption that the n macroeconomic series included in the VAR are all univariate *random walks with drift*.[8] Assuming the parameters to be uncorrelated with each other,

[8] The random walk hypothesis is imposed a priori. On the contrary, a posteriori each time series may follow a more complicated process if in the sample there is sufficent information to pull away from the prior.

the uncertainty around prior belief is represented by the main diagonal elements of the matrix Ω_0. The prior variance of the coefficients of the endogenous variables is specified as in equation (10). The so called "overall tightness" hyperparameter λ_2 describes the tightness of the prior distribution around the mean; it also describes the weight taken by sample information in the posterior distribution: when zero the prior is fully informative, i.e. the sample information is disregarded, while with an infinite value the prior is diffuse. The hyperparameter λ_3 controls the importance of the most recent lags relative to the more distant ones; the information content may decay according to a harmonic function of the lag order l.[9] This variance also depends on the terms $\widehat{\sigma}_i$ and $\widehat{\sigma}_j$, which represent the entries of the main diagonal of the covariance matrix of the measurement equation. As they are a priori unknown, following Litterman (1986), we consider the consistent estimates derived from the residual variances of simple AR(p) models with a constant term estimated through OLS. The ratio $\widehat{\sigma}_i/\widehat{\sigma}_j$ is included in the specification to correct for differences in the units of measurement of the endogenous variable or for other scale effects.

The hyperparameter $f(i,j)$ is known as the "relative tightness" and reflects the weight of the endogenous variable j in the equation for variable i. As the dependent variable's own lags are assumed to contain more information than the lags of the other endogenous variables, it is generally assumed to lie between zero and one, with a low value indicating weak interaction between variables i and j and a high value, strong interaction. A relevant question may be the choice between a symmetric and a general structure for the relative tightness. Many authors argue that using the general prior simply transforms the problem of over-parameterisation into having to estimate or search over too many hyperparameters; but others (such as Doan et al., 1984; Litterman, 1986; Amisano, Serati and Giannini, 1997) have shown that, especially when dealing with large systems (six or more endogenous variables), the choice of a symmetric prior may be inappropriate. In most of the BVAR models we opted for simplicity and used the symmetric prior,[10] namely, we imposed the

[9] Sometimes the information content may also decay according to the steeper geometric lag decay function, which implies λ_3^{l-1} with $\lambda_3 \leq 1$.

[10] During our experiments we also allowed for more general structures of the relative tightness. In general, improvements in forecasting accuracy of the BVAR models are marginal. In addition, results on

following structure

$$f(i,j) = \begin{Bmatrix} \lambda_4 & j \ endogenous, & \forall i,j = 1,...,n \\ \lambda_6 & j \ exogenous, & \forall i,j = 1,...,n \end{Bmatrix} \tag{12}$$

but, sometimes, in order to be able to model with sufficient generality some foreign variables that are included in the model as endogenous, but which can be made approximately exogenous with respect to some goal domestic ones. In this case exogeneity may be reached by tightening the prior of zero on the coefficients for the other endogenous variables in the equation for the foreign ones, which is the same as assuming a univariate autoregression on their dynamics.[11] The same tips apply when some financial variables (stock returns, exchange rates) are used together with macro variables, as we expect lags of other variables in the former to be less important than in the latter.

Finally, the prior variance of the constant term is specified according to equation (11), depending on the hyperparameter λ_5. A high value implies that hardly any prior information is available on the value the constant may take, while a value of zero implies that the knowledge is complete. When the specification of the BVAR model also includes dummies to detect outliers, a prior mean of zero is imposed for each and the prior variance is handled in each equation separately through the further set of hyperparameters.

It is important to note that a BVAR is a general model, which encompasses other specifications. A VAR corresponds to a BVAR with equal weights for hyperparameters of the variable's own lags and those of the other variables ($f(i,j) = \lambda_4 = \lambda_6 = 1$), no lag decay using the harmonic lag function ($\lambda_3 = 0$) and no weight for the prior ($\lambda_2 = +\infty$). A univariate autoregressive model is a BVAR with no other variable lags ($f(i,j) = \lambda_4 = \lambda_6 = 0$). These special cases are used as benchmarks both in specifying the optimal prior distribution and in evaluating forecasting performance.

All BVAR models considered in the next sections are specified in levels rather than in growth rates in order to apply the restrictions implied by the Minnesota prior assumptions.

marginal predictive content of financial spreads are robust to this choice.

[11] This approach may be preferred to one that treats the foreign variables as predetermined ones, as it does not need them for the construction of the ex-post forecasts for these variables, and thus, avoids hampering or complicating the comparisons between different competitive BVAR models.

As pointed out by Sims et al. (1990), stationarity of the series is unnecessary in BVAR models.[12] The specification of the optimal lag order is not a crucial issue in the BVAR framework, as the effect of more lags can be reduced or offset by the decay function used in the prior distribution. The number of lags chosen is in any case the smallest number necessary to preserve the white noise structure for the error terms of the VAR estimation. The indications from commonly used tests, such as the different information criteria (Akaike, Schwartz, Hannan and Quinn) and the modified likelihood ratio test suggested by Sims (1980), have also been considered.

5.2. The Choice of the Hyperparameters

A key issue in the specification of the BVAR models has been the choice of the hyperparameter values, namely the vector $\lambda = (\lambda_0, \lambda_1, \lambda_2, \lambda_3, \lambda_4, \lambda_5, \lambda_6)$. In a pure Bayesian approach they would reflect the beliefs of the researcher before analyzing the data, but in most of the BVAR literature, the data are also used to select the most appropriate values for the hyperparameters. In searching for optimal hyperparameter values, the choice of the objective function is crucial. As the main use of the BVAR models is for forecasting, many authors opted for a measure of forecasting accuracy. In this framework Doan, Litterman and Sims (1984) minimized the log-determinant of the h-step-ahead forecast error variance-covariance matrix, while Litterman (1986) and Amisano, Serati and Giannini (1997) opted for minimizing the Theil U statistic equation-by-equation at a given forecasting horizon. In a previous empirical investigation we estimated the hyperparameters, minimizing the Theil U statistic equation-by-equation four quarters ahead, as this horizon is considered of major interest for professional forecasters.

However, that approach led both to operational problems and to difficulties in the interpretation of the results. The first operational problem was choosing the training sample period for "tuning" the prior distribution through the ex-post forecasting performance. The BVAR literature suggests that this procedure sometimes may lead to an over-fitting problem, as the results may depend on the size of the "false in-sample forecast period"

[12] Engle and Yoo (1987) argued that in the case of cointegrated variables case the Minnesota prior may be inappropriate. However, Alvarez and Ballabriga (1994), with a Montecarlo experiment, found no significant improvement in modifying the prior to take an error correction term into account.

used for the prior calibration process: there is a trade-off between the need to have a robust prior distribution, obtained with a large sub-sample, and the risk of including too many structural breaks relative to relationships between variables, implied by a shorter one. Both problems were quite significant, as in the sample period there were major institutional and structural changes in the euro area: the oil shock at the beginning of the 1980s, German unification, the ERM-II turmoil at the beginning of the 1990s, the change in monetary regime for some euro-area countries and, finally, the move to an environment of price stability with the transition to Stage Three of EMU. In addition, we note that the choice of the forecasting horizon is somewhat arbitrary and not a trivial issue, as changing the horizon often led to different optimized hyperparameter configurations.

Following Canova (2002), who provided a natural answer to these operational problems, we estimated the vector λ by means of the maximization of the conditional predictive density

$$L(\lambda/y) = \int L\left(\beta/y, \lambda\right) \cdot g\left(\beta/\lambda\right) \cdot d\beta \tag{13}$$

which can be constructed and evaluated numerically in the entire sample period using the Kalman filter. Specifically, using the prediction error decomposition we can write the log-likelihood function for the VAR model as

$$\log L(y_1, ..., y_T) = \frac{T \cdot n}{2} \log\left(2\pi\right) - \frac{1}{2}\sum_{t=1}^{T} \log\left|\widehat{\Sigma}_t\right| - \frac{1}{2}\sum_{t=1}^{T} \widehat{\epsilon}_t \widehat{\Sigma}_t^{-1} \widehat{\epsilon}_t \tag{14}$$

where $\widehat{\epsilon}_t = y_t - \widehat{y}_{t/t-1} \sim N\left(0, \widehat{\Sigma}_t\right)$ is the one-step-ahead prediction error and where the initial conditions are such that $y_1 \sim N\left(\mu_1, \Sigma_1\right)$ and $\widehat{\epsilon}_1 = y_1 - \mu_1$. Prediction error decomposition is convenient in two respects. First, the building blocks of the decomposition are the forecasting error $\widehat{\epsilon}_t$ and their MSE $\widehat{\Sigma}_t$, so it can be used for for any model, either statistical or economic, that has an ARMA representation. Second, the hyperparameter estimates can be interpreted as minimising the one-step-ahead prediction error in the sample.

The search procedure was developed through the simplex method. This approach provides greater flexibility in the choice of the hyperparameters than in the default parameter settings available in the RATS routines, and in general gives better forecasting results.

The starting values of the hyperparameters were chosen using simple rules of thumb or previous experience. They are given by the vector $\lambda = (1.00, 10^{-7}, 0.2, 1.0, 0.5, 0.01, 0.5)$. When the maximizing algorithm failed to converge, the analysis was supplemented with a rough search over a grid of values for the hyperparameters.

5.3. Forecasting Strategy

The coefficients of the models and the hyperparameters were estimated through the Kalman filter over the 1980q1 to 1993q4 sub-period, leaving 1994q1-2002q4 for examining the accuracy of out-of-sample forecasts at horizons one to eight quarters ahead. This allows proper estimation while leaving a sufficient number of observations to assess the out-of-sample forecasting accuracy of the different models.

To exemplify, the models were estimated through 1993q4, and a set of out-of-sample forecasts 1 through 8 steps ahead were computed, spanning the period 1994q1-1995q4. Next, the 1994q1 observation was added to the estimation sample, and a new set of forecasts, from 1994q2 to 1996q1, was computed. The process was iterated up to the end of the available dataset. The forecasting properties of the models were assessed using the resulting collection of 1 through 8 step ahead forecasting errors.[13] Specifically, we collected the root mean square error (RMSE) of each model considered and computed the Theil U statistic - the ratio of the RMSE of the model forecasts to the RMSE from a random walk model- at each forecasting horizon.

The paper distinguishes the different indicators according to the timing of the release, in order not to bias the results in favour of indicators (e.g. real GDP) that tend to be available much later than others (e.g. financial variables) that are available with no delay.[14] Following Doan, Litterman and Sims (1984) and Amisano, Serati and Giannini (1997), the procedure used to handle this problem is the conditional forecasting approach,

[13] For the Bayesian models the hyperparameters were not re-estimated each time; they were kept constant at the values obtained for the initial training sample period (1980q1-1993q4).

[14] The entire data-set used for estimation and forecasting is based on revised - as opposed to real-time - time series. This might in principle bias the results against finding significant forecasting power in financial spreads, as non-financial variables are typically those that most benefit from subsequent revisions. However, putting together a real-time data-set, difficult under normal circumstances, is virtually impossible in the EMU context.

which looks at the orthogonalized vector moving average representation of the BVAR and is based on the idea that conditioning on future values of the endogenous variables vector entails constraining some future values of the orthogonalized disturbances to be non-zero.[15]

To assess the predictive content of financial spreads we opted for a two-stage strategy. First, we develop a BVAR model for the euro area with good forecasting performance, to be used as a benchmark. Then we add the spreads to this benchmark, one at a time, and gauge the gain or loss in forecasting accuracy.

5.4. The Benchmark Model

The baseline model comprises real GDP, inflation, the short-term interest rates and a measure of the money supply. M1, M2 and M3 are considered. All models also include a constant term. Table 4 and 5 show the Theil U statistics for these models at horizons from one to eight quarters ahead.

For comparison, we also report results for alternative models suggested in the literature. Specifically, two univariate specifications were considered: Box-Jenkins models and time varying coefficients Bayesian AR models (TV-AR). The specification search of the ARIMAs was conducted based on the usual identification approaches, as we looked at the stationarity, autocorrelation and partial autocorrelation functions, significance of coefficients, and the Akaike Information Criterion for the choice of lags. The identification procedure suggested an ARIMA (1,1,0) for real GDP and an ARIMA (3,1,0) for inflation. On the contrary, the TV-AR models are specified with five lags, allowing the prior distribution eventually to tight or loose the more distant lags.

Simple ARIMA models generally perform better than the TV-ARs at all the horizons considered, while the BVAR models strongly dominate univariate specifications especially at horizons longer than 1 or 2 quarters. We notice that the gains deriving from multivariate specifications are larger for the inflation rate. For real GDP forecasts, the inclusion of M2 leads to better performance up to six quarters ahead. Surprisingly, the model with M1

[15] More details are given in Doan, Litterman and Sims (1984) and in the RATS 5.0 Reference Manual.

seems to perform worse both at shorter and longer horizons. The latter result is not in line with recent empirical work of Agresti and Mojon (2001) and Brand (2000) for the euro area. For inflation predictions, the inclusion of M2 and M3 leads to good forecasting performance relative to the model with M1. In particular, both models yield better forecasts at all horizons. The model comprising M2 outperforms that with M3 at horizons from three to seven quarters ahead, while at shorter horizons the model with M3 does better. Some models comprising two monetary aggregates simultaneously were also explored, but the results were not encouraging. Altogether, considering that the model comprising M2 does best at horizons of 1 year, which is often crucial in forecasting exercises, we opted to include the latter aggregate in the benchmark model; in the following subsection, however, we check the robustness of the results by experimenting with alternative benchmark models comprising both M1 and M3. For each of the multivariate specifications we also set up unrestricted VAR models including the same endogenous variables, but without a Minnesota prior. We found that they are not very successful in forecasting real GDP and inflation because of the large number of parameters to be estimated. This is noticeable in particular for real GDP forecasts at medium term horizons, where the Theil U statistics are significantly worse than those obtained with both univariate specifications and Bayesian VAR models. Inflation forecasts with VARs are better up to two quarters ahead. At longer horizons forecasting accuracy deteriorates sharply and BVARs dominate.

In a second step, these domestic models have been extended to take into account the possible role of some international variables. The inclusion of the real or nominal effective exchange rate of the euro does not help forecast either target variable at any horizon. The inclusion of the total and the non-energy commodity price indexes, or the price of oil does not improve inflation predictions, although it helps forecast GDP marginally at shorter horizons (up to four quarters ahead). Better results are obtained using some US macroeconomic variables: both real GDP and short-term interest rates improve euro-area real GDP forecasts at longer horizons (from five to eight quarters ahead); the US real GDP performs best. No improvements at the same horizons are obtained in forecasting euro-area inflation.

Based on these results, the following five variables were included in the benchmark BVAR model: real GDP, consumer inflation, the short-term interest rate, the money aggregate M2 and the US real GDP. The last row in tables 4 and 5 presents the results obtained with this benchmark BVAR, allowing a comparison with both the domestic models and the univariate specifications.

For completeness, in Table 6 we report hyperparameter estimates for all the Bayesian models considered in the analysis. We argue that adding time variation to the coefficients substantially improves forecasting performance. Potential gains in accuracy are already evident with univariate specifications. A time-varying AR model produces better U Theil statistics than a fixed coefficient ARIMA model for both real GDP and inflation. The picture is the same when multivariate models are considered. Anyway, the ex-post validations seem to suggest that the forecasting accuracy of the euro-area BVAR models is very sensitive especially to the calibration of the time-varying hyperparameter. Even if a fixed parameter model does not appear suited to deal with a non-stationary environment (its forecasting performance is drastically worse than that of a time-varying model), the optimal amount of time variation selected by the procedure is generally small, but even small positive deviations from the optimal value produce significant changes in forecasting performance.

In the specification of the benchmark model there is another relevant issue. The derivation of the Kalman filter assumes that the innovations in the measurement and observation equations are both normally distributed. In the context of a constant coefficient state space model the misspecification of the distribution of the errors does not create consistency problems, as the maximum likelihood estimates obtained incorrectly assuming a normal distribution (typically called quasi-ML) have nice properties under a set of regularity conditions.[16] On the contrary, when the coefficients of the state space model are time-varying, checking normality of the residuals is a relevant issue, as ML estimates obtained with misspecified errors are no longer asymptotically equivalent to

[16] We could reached the same conclusion by noting that recursive OLS estimates are consistent and asymptotically normal if the regressors are stationary, ergodic and uncorrelated with the error and that recursive OLS and Kalman filter estimates coincide if a conditional likelihood is used.

those of the correct model and Kalman filter estimates cannot be interpreted as the best linear estimates of the coefficient vector minimizing the mean square error. To this end, we report in Table 7 summary statistics on normality of estimated residuals for each equation of the benchmark model. For euro-area variables we essentially notice absence of any misspecification. On the contrary, the distribution of the estimated residuals for the US real GDP equation seem to be significantly skewed. Looking at the estimated residuals derived from the Kalman filter coefficient estimates (Fig. 3), we argue that the violation of normality assumptions is due to the presence of some significant outliers concentrated in first part of the sample period.[17]

5.5. The Marginal Predictive Content of the Spreads

Next, the different measures of the financial spreads described in section 3 were added in turn to the benchmark model, producing a new set of forecasts for the target variables, using the same methodology.

Tables 8 and 9 report the Theil U statistics of the models with each spread in order to assess gains or losses in forecasting accuracy with respect to the benchmark model. The first thing to note is that these gains, where they exist, are minuscule. The inclusion of the slope of the yield curve improves the inflation forecasts at horizons from four to eight quarters ahead, but the gains are very small (the Theil U statistics are lower by around 4 per cent on average). For real GDP, the marginal predictive content of the slope of the yield curve seems to be zero. We notice that real GDP forecasts worsen at all horizons considered, and the losses are larger as the forecasting horizon lenghtens.

Better results may be obtained for inflation as well as for real GDP using the long-term interest rate rather than the slope of the yield curve. The inflation forecasts are better than those of the benchmark model at all horizons considered; the gains in accuracy are again greater four-to-seven quarters ahead by around 7 per cent. The maximum gain (around

[17] During our analysis we also set up BVAR models including dummy variables for controlling some large outlier in the euro area equations. Specifically, we considered a dummy controlling for the German unification (1991q1) and the beginning of the EU (1993q1). When M1 was considered, the model included an additional dummy controlling for its exceptionally large increase reflecting the transition to Stage Three of EMU in 1999q1. Improvements in forecasting accuracy were negligible.

10 per cent) is at four quarters ahead. For real GDP predictions, there is no improvement over the benchmark model, but the losses in accuracy are smaller than those obtained under the model using the slope of the yield curve. Even if the same information set is used, different estimation techniques may lead to different results due to the specification of the prior distribution, which is included in the Kalman filter estimation through the initial state vector and the initial covariance matrix of the states.

The model comprising the reverse yield gap, measured either by earnings yields or dividend yields, does not outperform the inflation forecasts obtained with the benchmark model at any horizon. However, the inclusion of the reverse yield gap based on earnings yields seems to lead to some improvements in real GDP forecasts one-to-two quarters ahead, although the gains in forecasting accuracy are tiny (around 2 per cent). Finally, the credit quality spread does not improve either the real GDP or the inflation forecast at any horizon.

We also check whether results on the marginal predictive content of spreads are robust to the choice of the sample period, by repeating the exercise for the sub-sample 1999q1 and 2002q4. The picture is essentially the same for all the spreads considered. The inclusion of the slope of the yield curve improves the inflation predictions two to eight quarters ahead, with the maximum gain, around 2 per cent, reached at four to five quarters. Again, at all horizons the inclusion of the long-term interest rate outperforms the benchmark model (by around 10 per cent on average) and the model comprising the slope of the yield curve (by around 8 per cent on average). The model comprising the reverse yield gap, based on earnings, does not improve real GDP forecasts at any horizon. The credit quality spread is the only one to show some marginal predictive content for real GDP at shorter horizons (from one to three quarters ahead), although less than 1 per cent on average.

Another way to assess the marginal predictive content of financial spreads is to evaluate their ability to predict turning points. To this end, Figures 6-13 show real output growth and inflation forecasts obtained with each model comprising a different spread, allowing a comparison with both historical values of target variables and forecasts obtained with the benchmark model. Financial spreads do not appear to have any additional predictive content. None of the spreads is able to track the direction of changes in real GDP growth and inflation rates. The peaks and troughs of the predicted business cycle are essentially the same as in the benchmark model, with no improvement in the ability to capture turning points. For real GDP growth rates, all the models capture the troughs in 1996q1, 1998q4 and 2001q1, and the peak in 2001q3 with a 1-quarter lag. For inflation, the models forecast the troughs in 1998q4 and 2001q2 with a 1-quarter lag.

We also explored the robustness of the results with respect to the alternative benchmark model including the M1 and M3 monetary aggregates. The general pattern of results with the financial spreads is remarkably similar to that obtained for the model including M2. When the slope of the yield curve is included, inflation forecasts improve. When the reverse yield gap based on earnings is considered, real GDP forecasts are better at longer horizons. For these models, improvements are around the same magnitude as those obtained with the benchmark including M2. Hence, results on the marginal predictive content of financial spreads seem to be robust to the choice of the monetary aggregate.

6. Concluding Remarks

This paper deals with the usefulness of a number of measures of financial spreads (the slope of the yield curve, the reverse yield gap, the credit quality spread) for forecasting real economic activity and inflation in the euro area. A preliminary investigation on the information content of each financial spread is conducted by means of a cross-correlation analysis. The analysis shows that financial spreads are characterized by good leading properties, especially for real GDP, at horizons from four to eight quarters ahead, but that correlations with economic activity and inflation are generally not stable over time. Thus, financial spreads still represent pieces of useful information that can help guide European monetary policy.

In a second step, a quarterly Bayesian vector autoregression model is used to assess the marginal forecasting power of the spreads for real economic activity and inflation. A benchmark BVAR is set up, containing real GDP, inflation, and key indicators of monetary policy and foreign macroeconomic variables. The leading indicator properties of the spreads are then assessed by augmenting the benchmark BVAR with the spreads, one at a time and looking at the out-of-sample forecasting performance. To avoid potential instability of the predictive equations involving spreads and target variables, we use time-varying coefficient models, with particular attention to the information set so as not to bias the results in favour of indicators that are available later than others. We find that financial spreads do not appear to contain marginal predictive content for future output growth or inflation. Our thesis is that the inclusion of some monetary policy variables, such as short-term interest rates and monetary aggregates, as well as other potential leading indicators, crowd out the spreads. These features remain robust to the choice of monetary aggregate, as well as to different sample periods and forecasting horizons.

Overall, the results suggest that there is no ready-to-use financial indicator that can replace an encompassing multivariate model for the prediction of inflation or output growth in the euro area.

Appendix I. Data sources and construction of the variables

Real GDP. Official real GDP is available from Eurostat only starting in 1991, as Germany is the limiting Member State, whose series on GDP on the basis of ESA95 starts only with the quarter after unification. Nevertheless, GDP data at single-country level exist for several Member states, including Germany, on the basis of ESA79, so an artificial longer time series for GDP may be compiled using them. As no uncontroversial aggregation method exists, we take the series of the ECB's area-wide model (AWM) by Fagan, Henry and Mestre (2001), aggregated on the basis of GDP weights at PPP exchange rates of 1995. Hence, the official time series from Eurostat has been backdated with the AWM's quarter-over-quarter growth rates.

Inflation. Euro-area inflation is measured by the quarterly averages of monthly data of the Harmonized Consumer Price Index (HCPI), the variable chosen by the ECB for the quantitative definition of its primary objective of price stability. The monthly series for the aggregate euro area has been officially collected only since in 1995, but it has been extended, backward up to January 1980 by Eurostat, aggregating individual countries seasonally unadjusted CPI indexes with consumption expenditure weights at the irrevocably fixed conversion rates of 31 December 1998. The reconstructed index has been seasonally adjusted with TRAMO-SEATS.

Monetary aggregates. The monetary aggregates (M1, M2 and M3) have been taken from the database built by the ECB to construct the historical time series of monetary aggregates for the euro area. The series are quarterly averages of the end-month stocks of M1, M2 and M3, seasonally adjusted.

Interest rates. The short-term interest rates are three-month money market rates; long-term interest rates are 10-year benchmark government bond yields or close substitutes. The EU-11 aggregates are taken from the AWM database. After 1998q4, they are updated, respectively, with the three-month and the Euro 10 year benchmarks as published in the ECB monthly bulletin. The slope of the yield curve is defined as the difference between the long-term and the short-term interest rates. For the construction of the credit quality spread we used lending rates on short-term loans to firms. The EU-11 aggregates are taken from the AWM database. After 1998q4 they are updated with time series published in the ECB monthly bulletin.

Stock market variables. The two versions of the reverse yield gap are defined as the difference between the long-term interest rate and, respectively, dividend yields and the earnings yields. The dividend yields and earnings yields ratios refer to the stock market Global Index from Datastream (TOTMKEM). It does not include all companies in the euro area stock market, but the most important ones according to market capitalization. As a caveat, it should be noted that the resulting stock market series does not use the same aggregation scheme as the other macroeconomic variables (euro-area output, inflation, monetary series and interest rates).

International variables. Total and non-energy commodity price index, as well as oil price time series, are taken from the IMF database. The real and the nominal effective exchange rate of the euro are taken from Bank of Italy. The US real GDP is billions of US dollars at 1996 prices (seasonally adjusted). The US short-term interest rates are the Federal Funds rates. They are taken from the database of the Federal Reserve Bank of St. Louis.

Figures and Tables

Figure 1

FINANCIAL SPREADS AND TARGET VARIABLES (1980:1-2002:4)

Figure 2

BUSINESS CYCLE COMPONENT OF FINANCIAL SPREADS

(1980:1-2002:4)

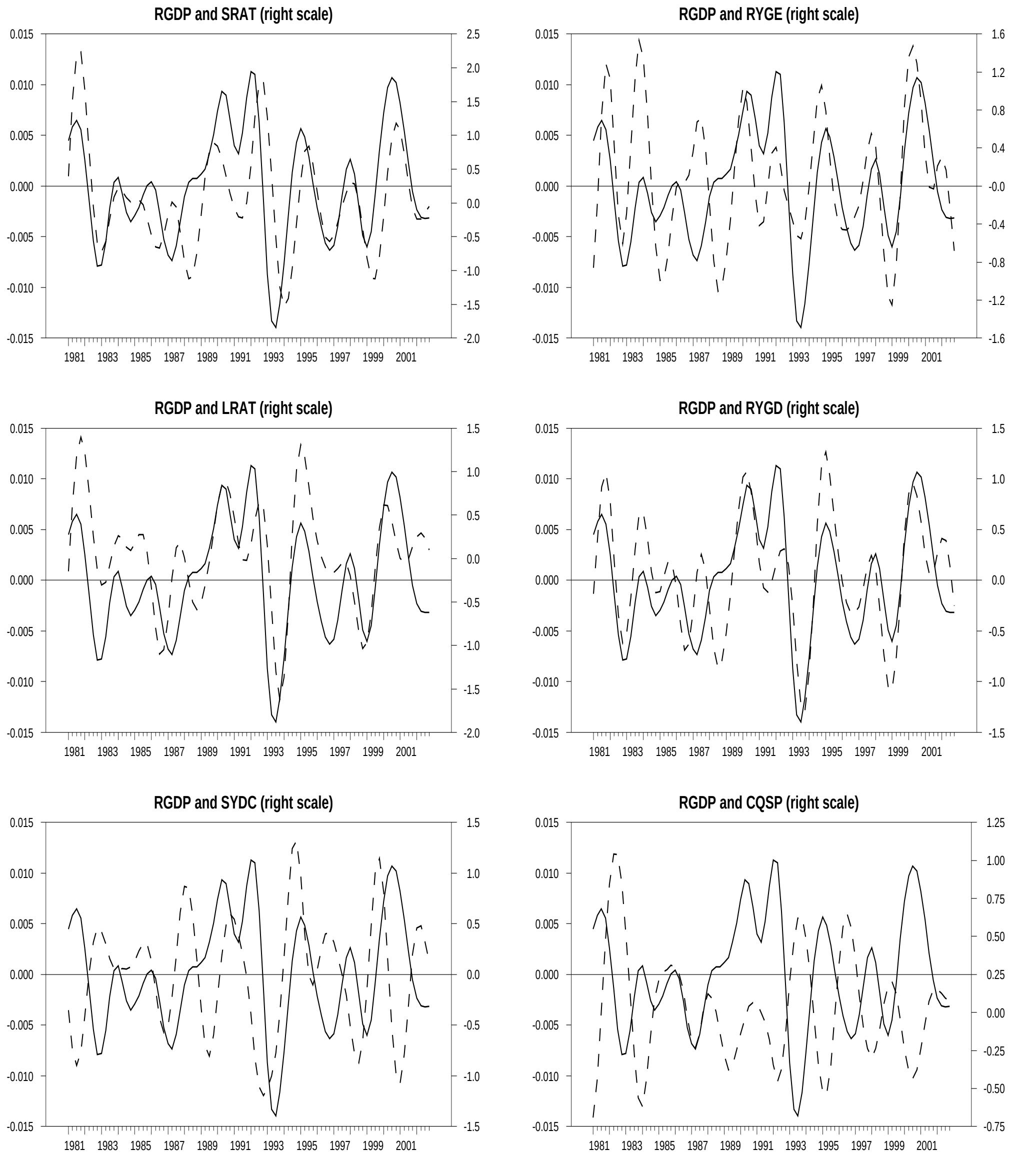

Figure 3

BENCHMARK MODEL: ESTIMATED RESIDUALS

(1982:2 - 2002:4)

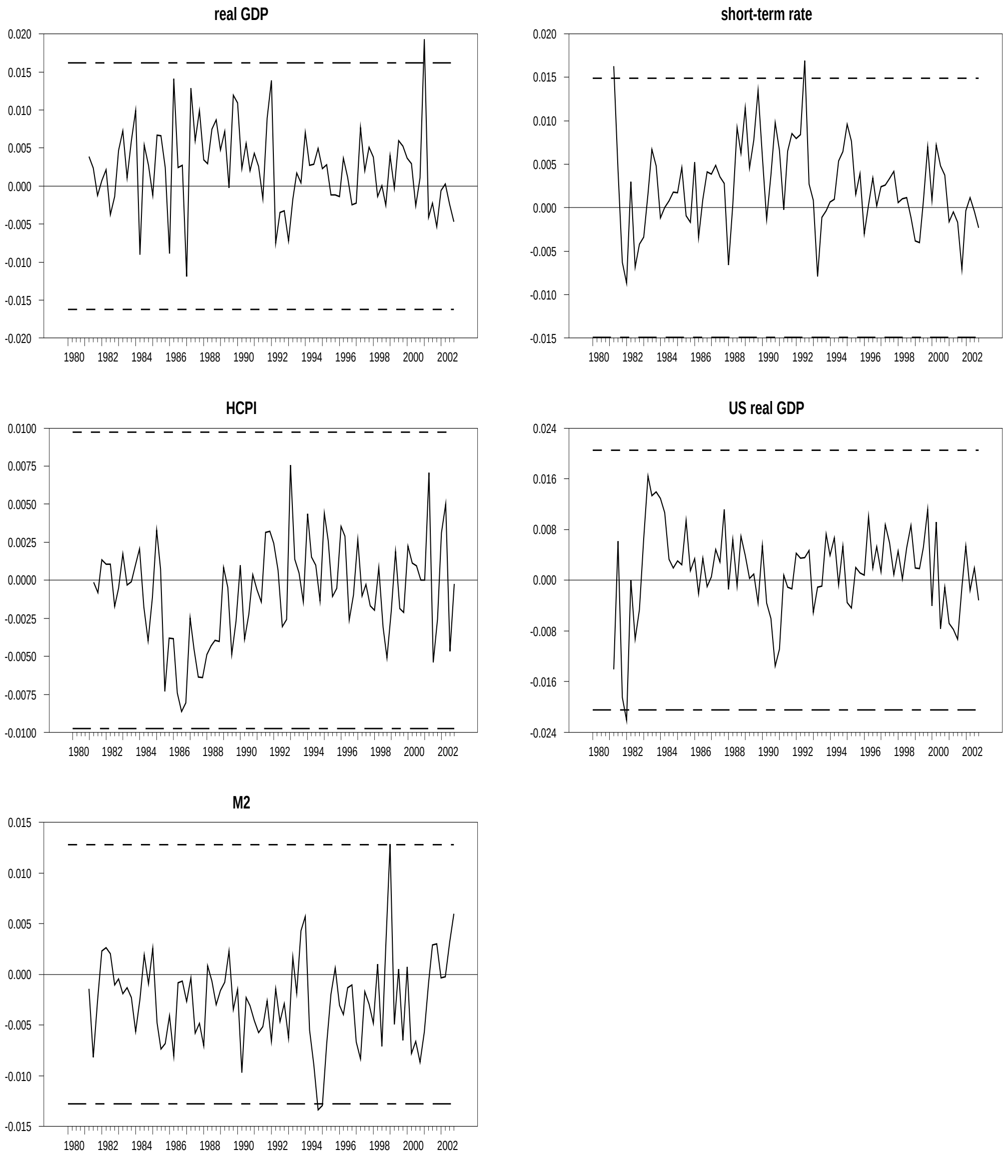

Figure 4

BENCHMARK MODEL - REAL GDP FORECASTS

(solid line: hystorical values; dashed line: forecasts)

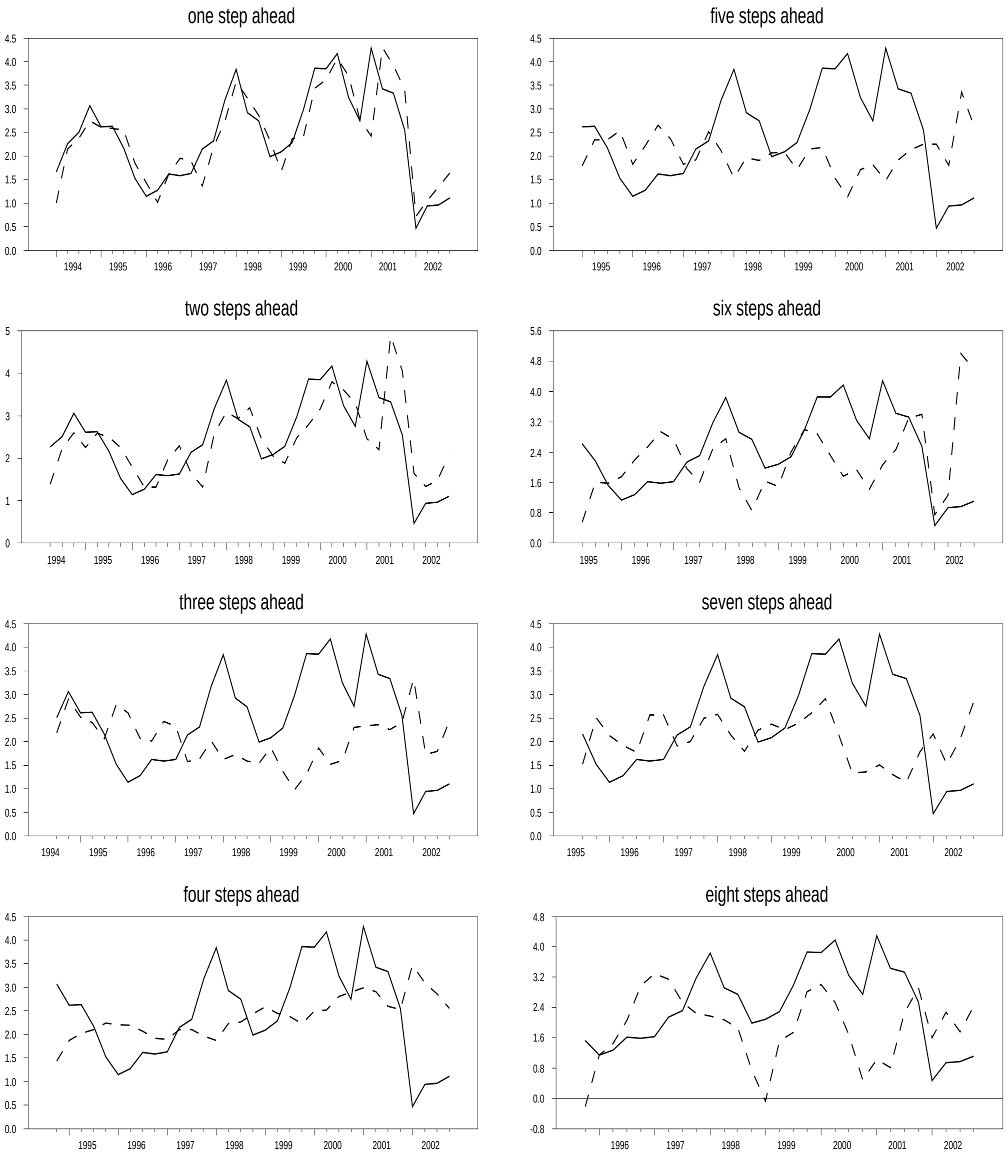

Figure 5

BENCHMARK MODEL - INFLATION FORECASTS

(solid line: hystorical values; dashed line: forecasts)

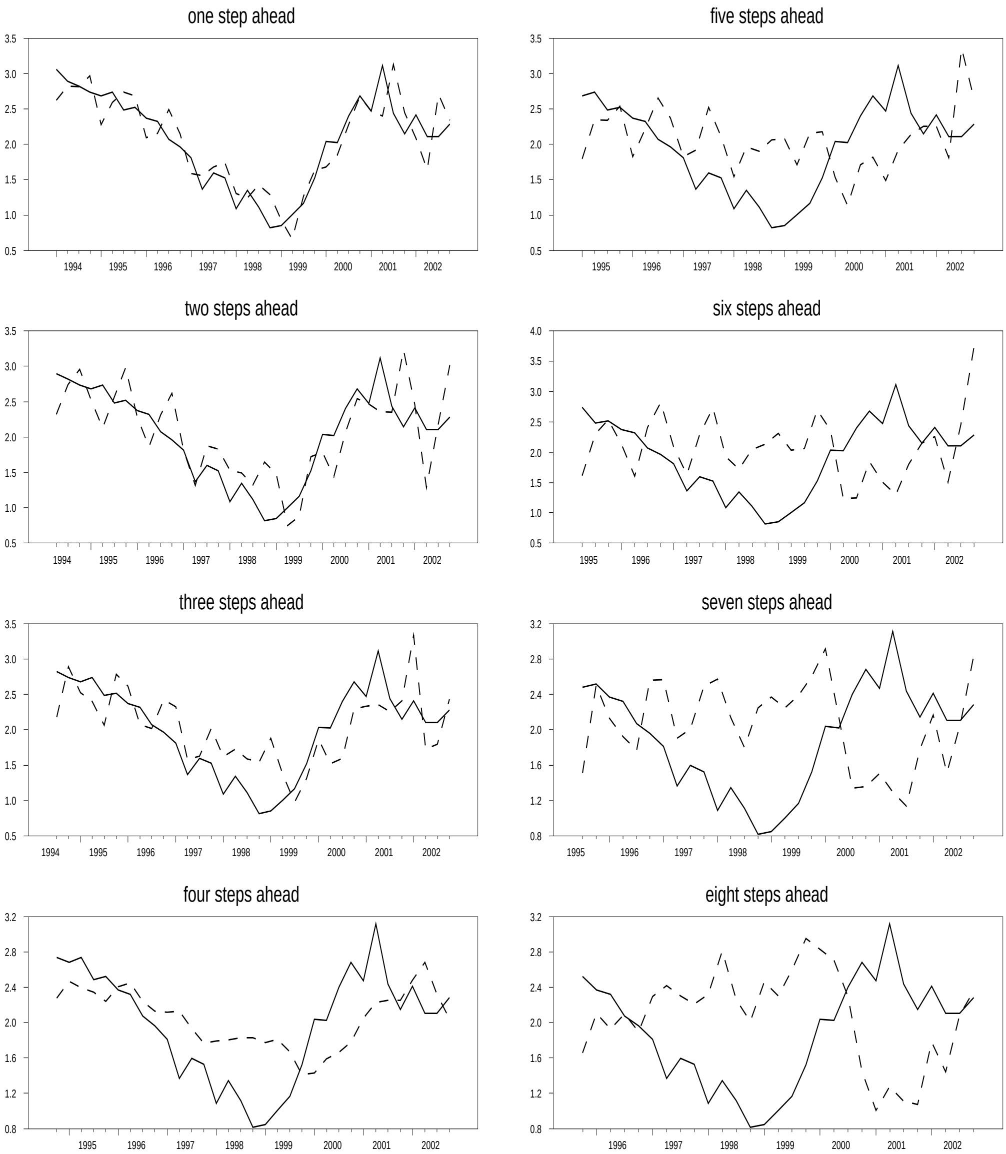

Figure 6

SLOPE OF THE YIELD CURVE - REAL GDP FORECASTS

(solid line: hystorical values; dashed line: forecasts)

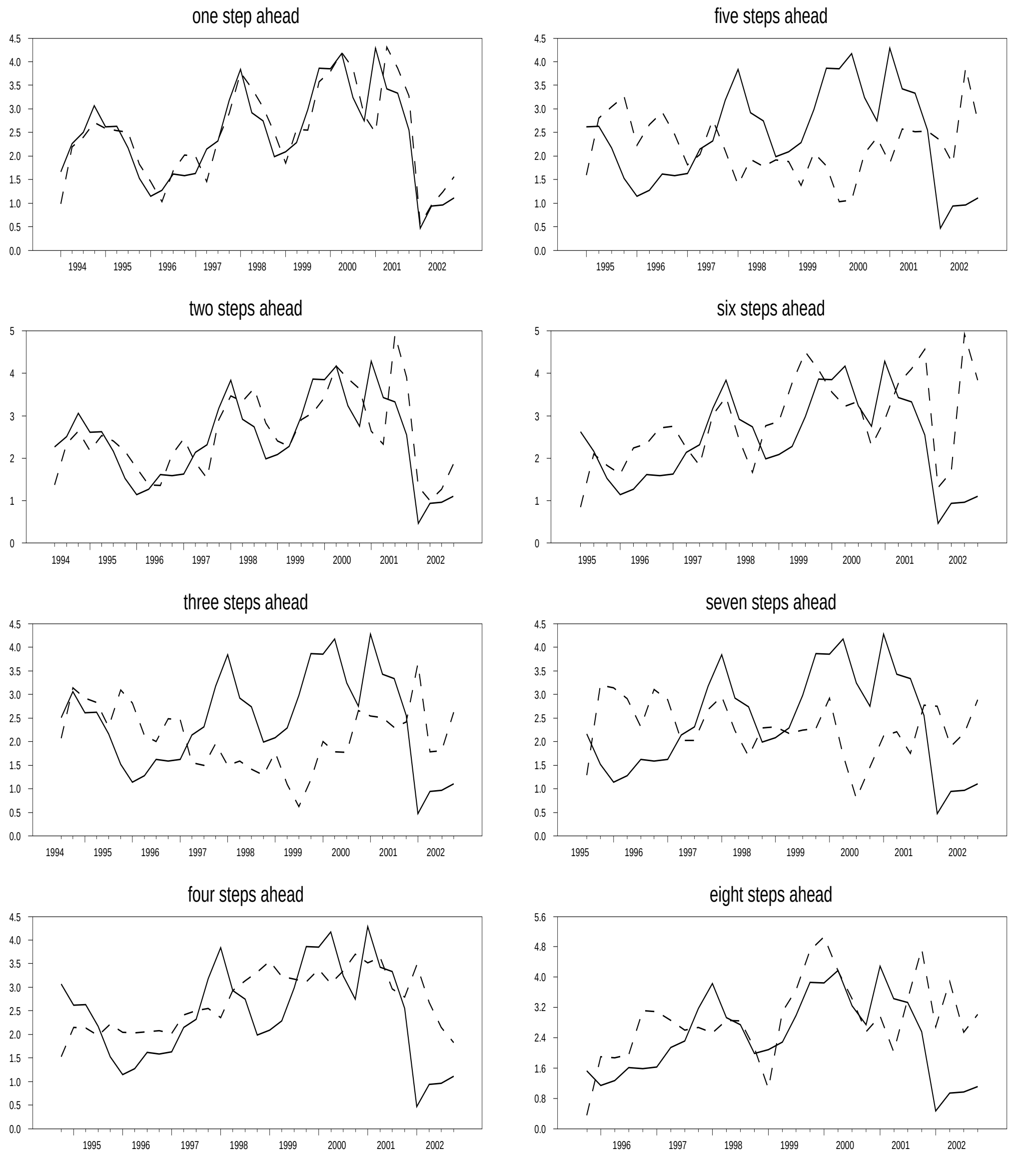

Figure 7

SLOPE OF THE YIELD CURVE - INFLATION FORECASTS

(solid line: hystorical values; dashed line: forecasts)

Figure 8

REVERSE YIELD GAP (DIVIDENDS) - REAL GDP FORECASTS

(solid line: hystorical values; dashed line: forecasts)

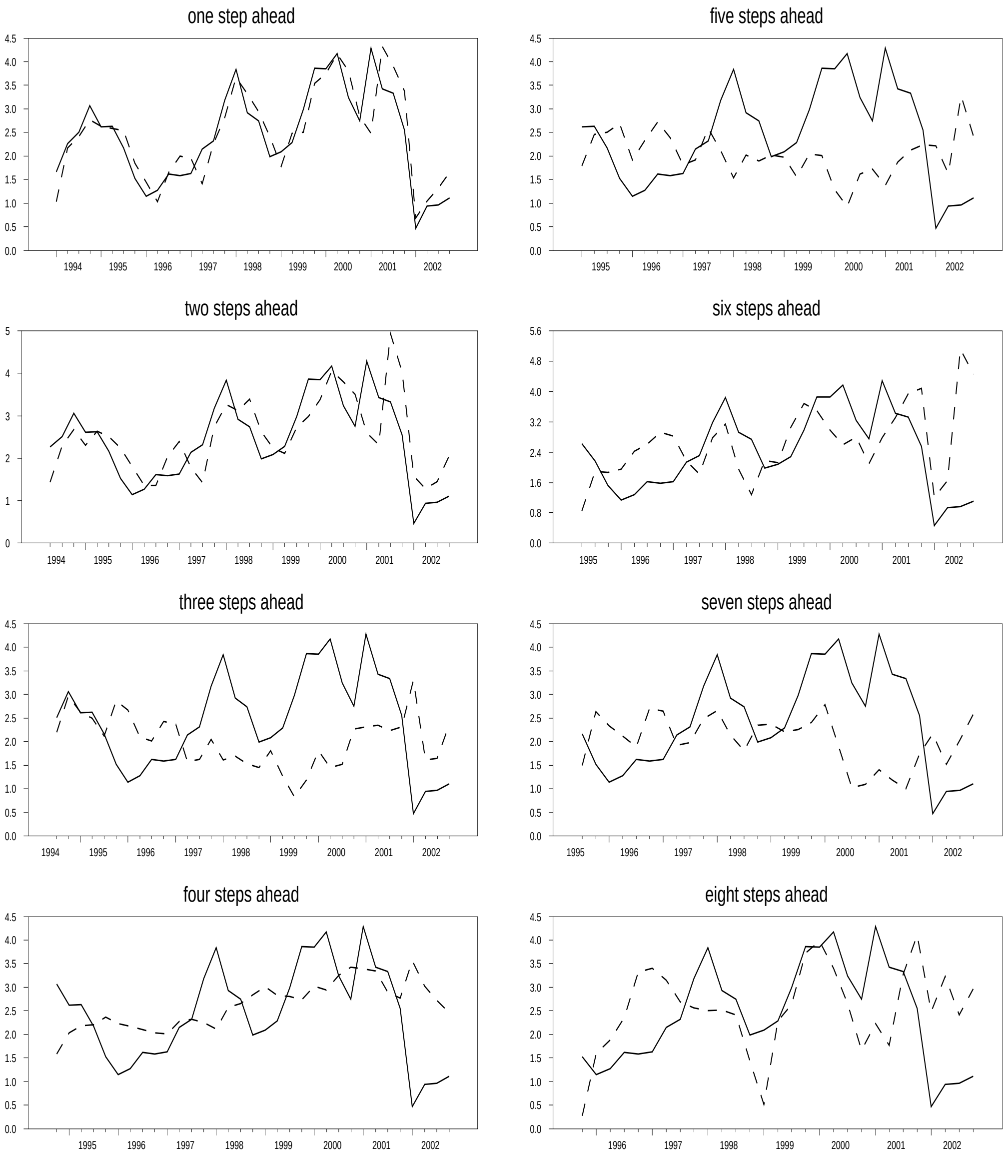

Figure 9

REVERSE YIELD GAP (DIVIDENDS) - INFLATION FORECASTS

(solid line: hystorical values; dashed line: forecasts)

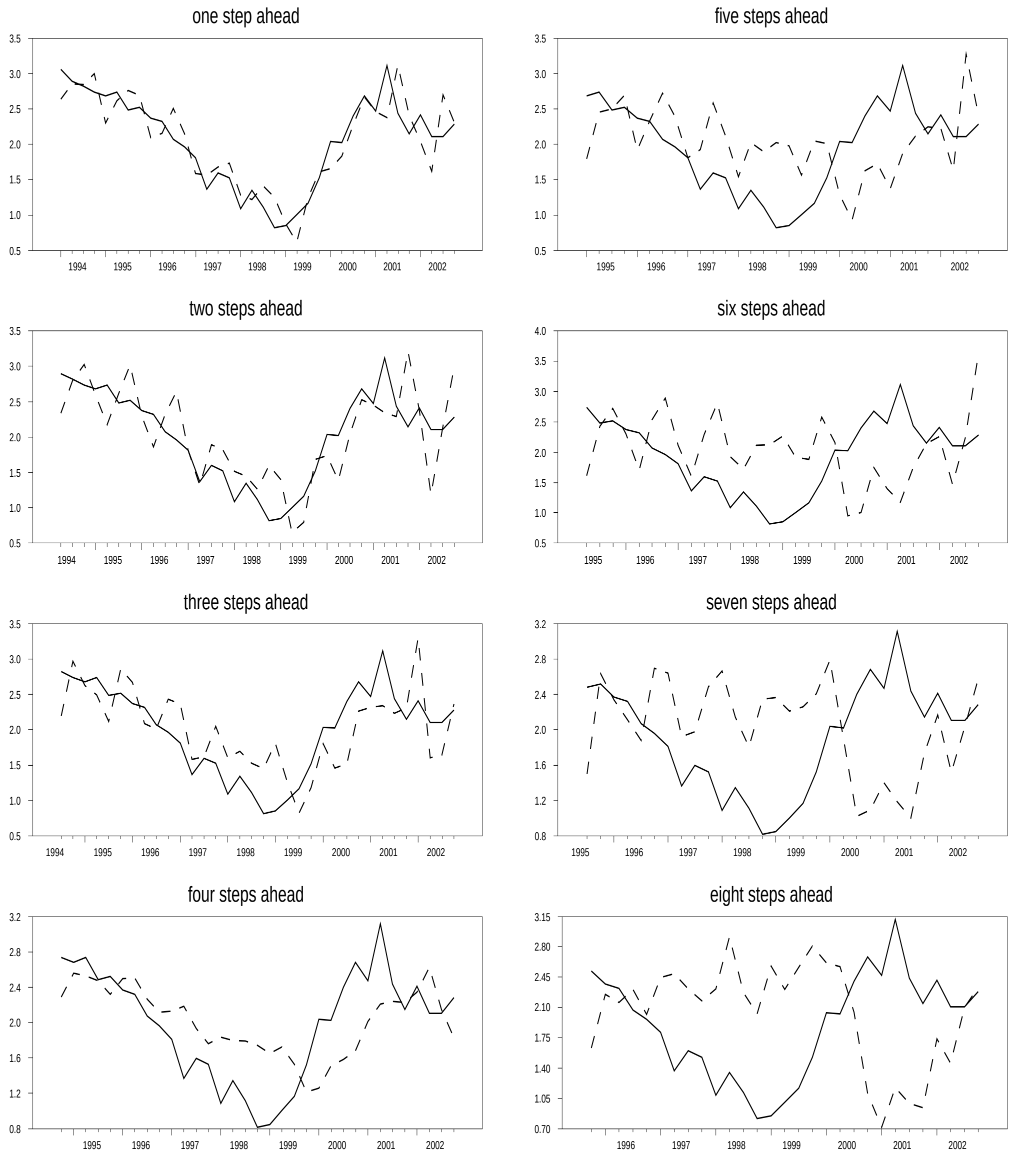

Figure 10

REVERSE YIELD GAP (EARNINGS) - REAL GDP FORECASTS

(solid line: hystorical values; dashed line: forecasts)

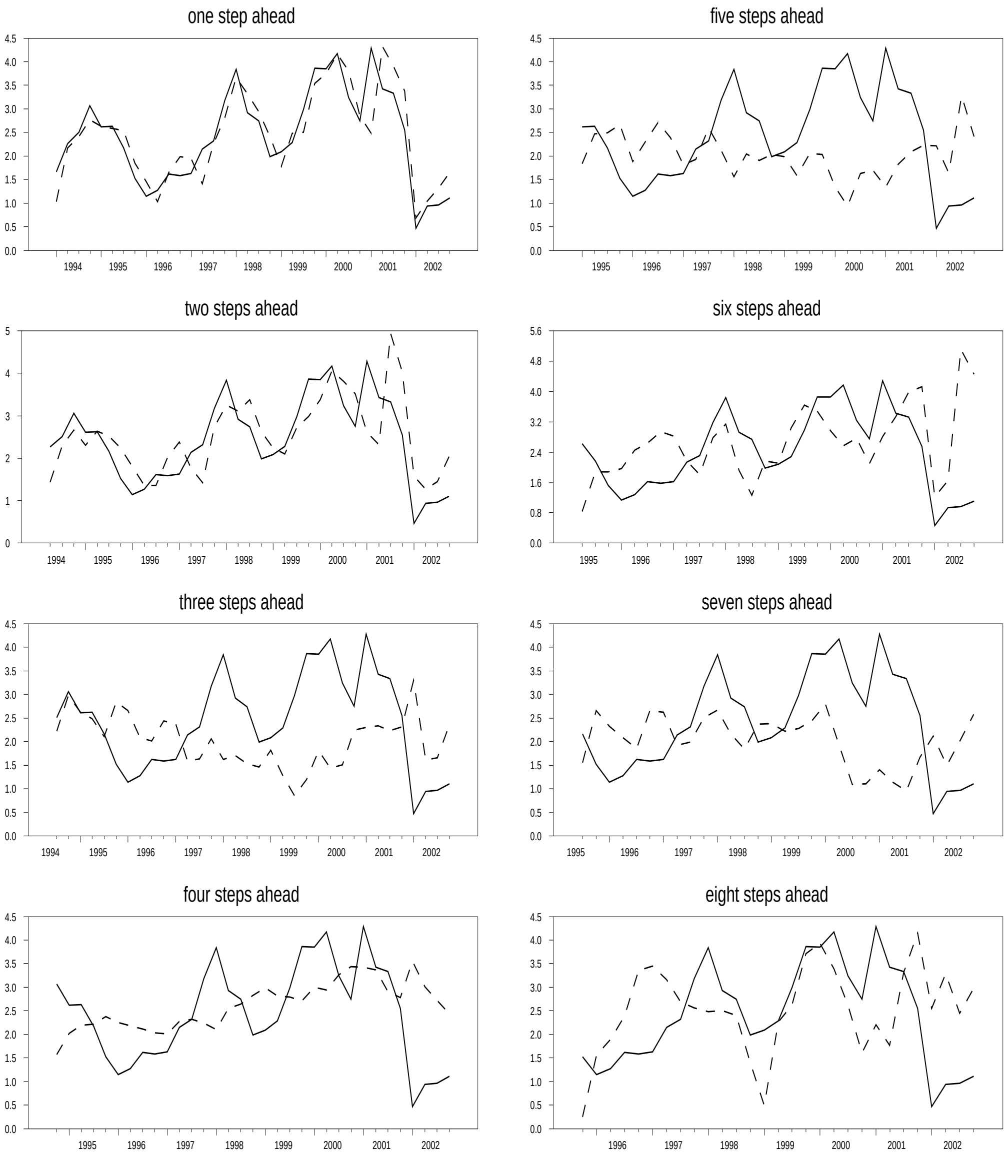

Figure 11

REVERSE YIELD GAP (EARNINGS) - INFLATION FORECASTS

(solid line: hystorical values; dashed line: forecasts)

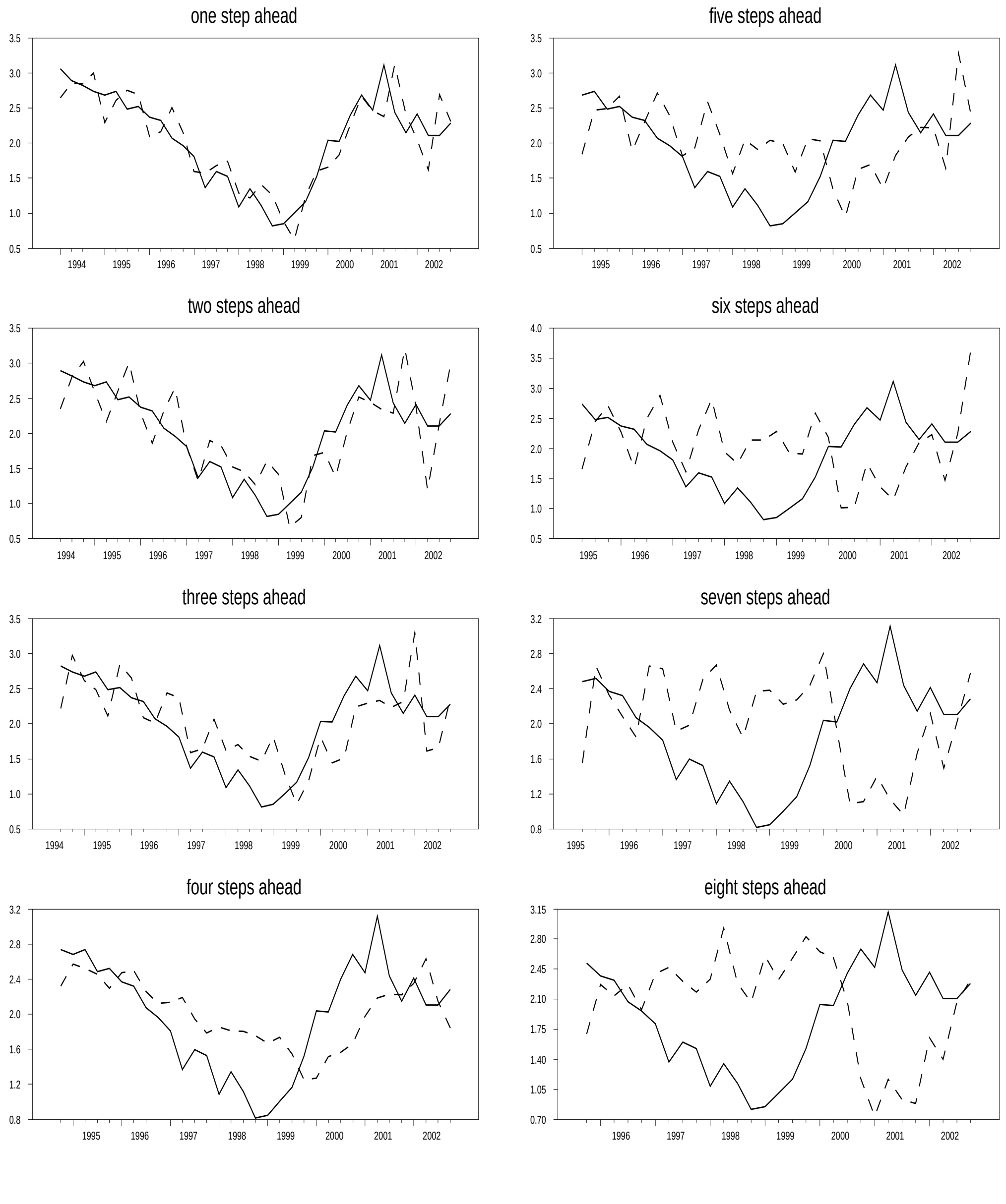

Figure 12

CREDIT QUALITY SPREAD - REAL GDP FORECASTS

(solid line: hystorical values; dashed line: forecasts)

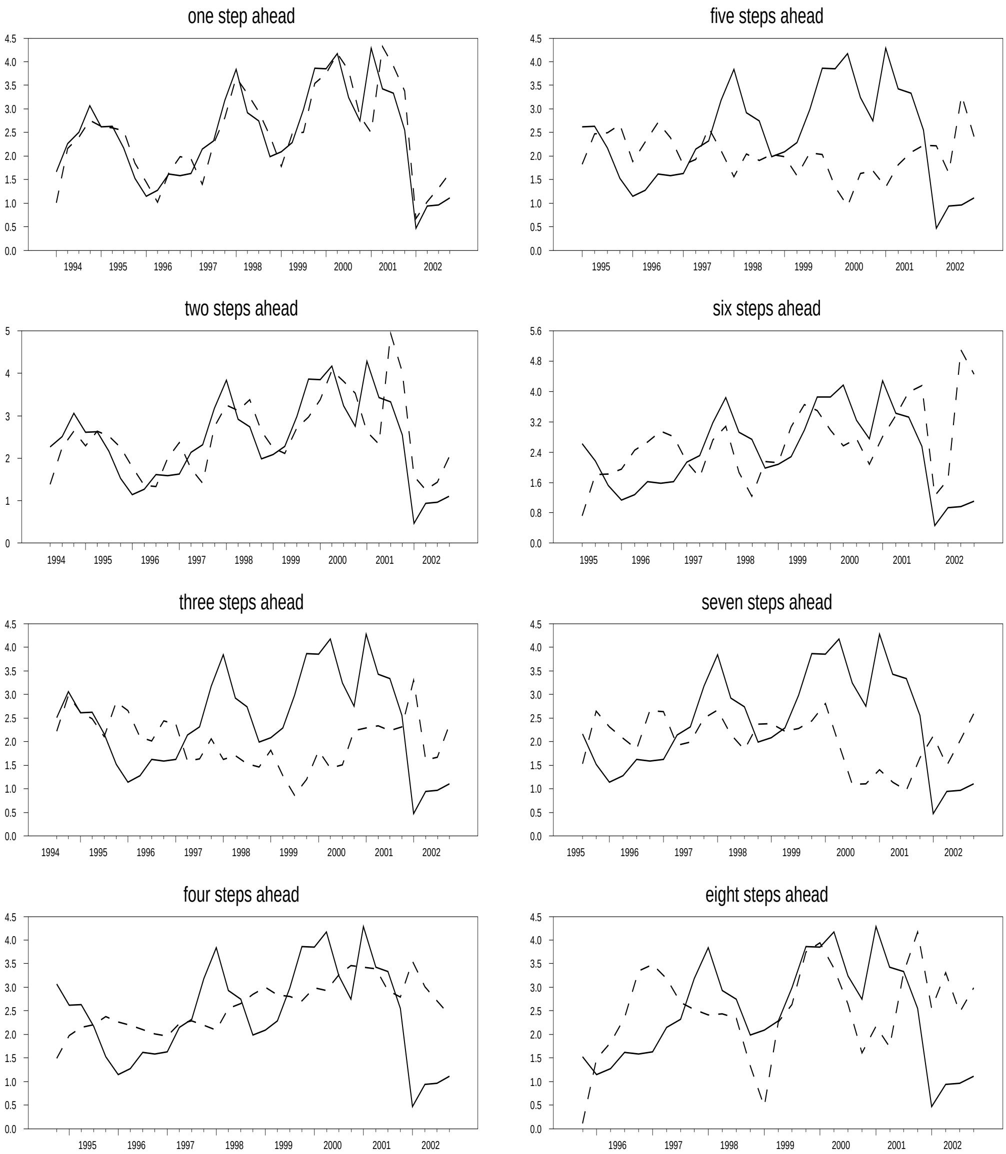

Figure 13

CREDIT QUALITY SPREAD - INFLATION FORECASTS

(solid line: hystorical values; dashed line: forecasts)

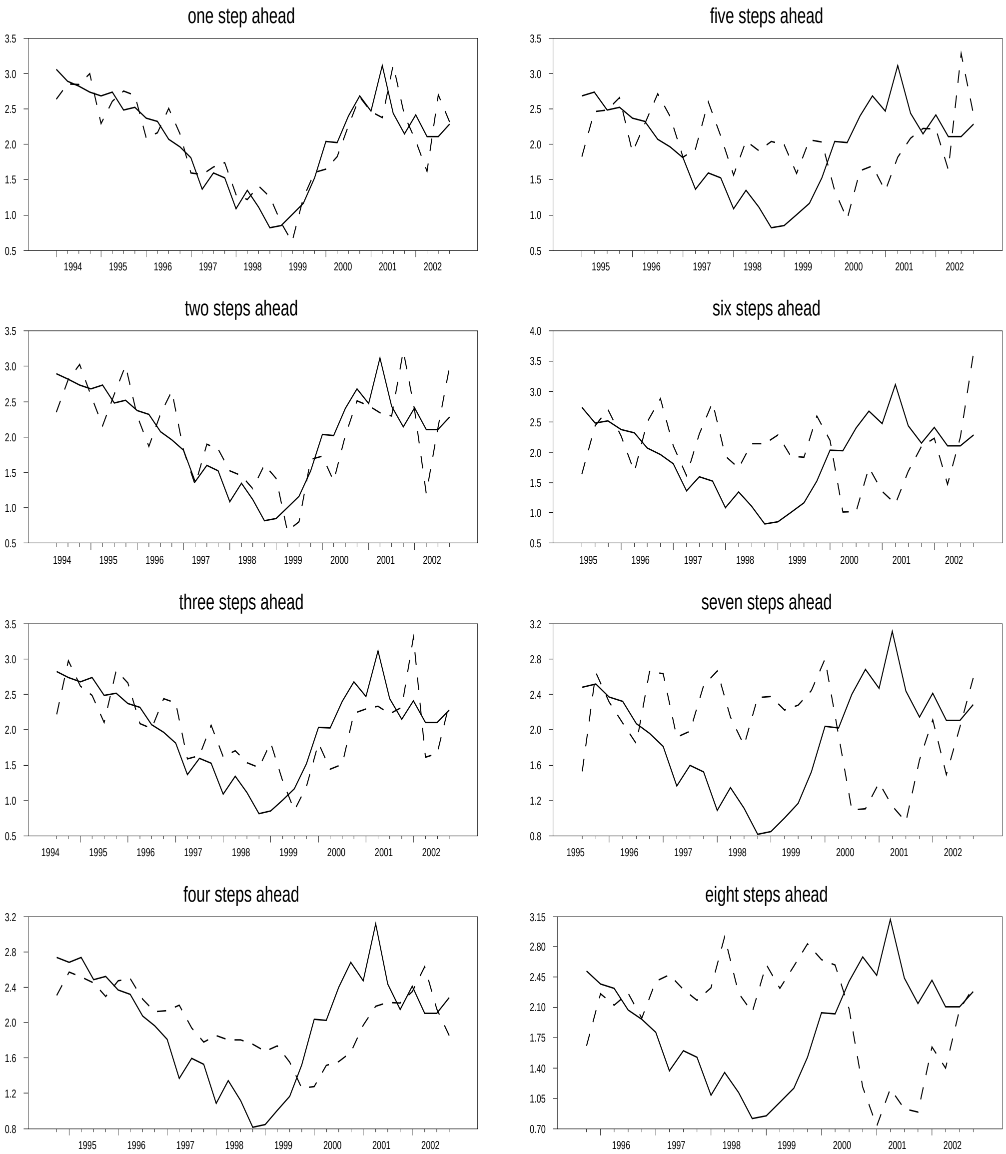

Table 1

VOLATILITY ANALYSIS OF FINANCIAL SPREADS

	1980:1-2002:4		1994:1-2002:4		1999:1-2002:4	
variable	absolute	% of RGDP	absolute	% of RGDP	absolute	% of RGDP
RGDP	0.56	1.00	0.50	1.00	0.58	1.00
HCPI	0.50	0.89	0.28	0.56	0.28	0.48
SRAT	0.78	1.39	0.68	1.36	0.74	1.28
LRAT	0.60	1.07	0.59	1.18	0.41	0.71
SYDC	0.59	1.05	0.63	1.26	0.68	1.17
RYGD	0.58	1.04	0.60	1.20	0.56	0.97
RYGE	0.67	1.20	0.68	1.36	0.79	1.36
CQSP	0.37	0.66	0.31	0.62	0.22	0.38

NOTE: RGDP is real gross domestic product, HCPI is the harmonized consumer price index; SRAT is the short-term interest rate; LRAT is the long-term interest rate; SYDC is the slope of the yield curve; RYGD and RYGE are two measures of the reverse yield gap, based, respectively, on dividend yields and earning yields; CQSP is the credit quality spread. The columns report, respectively, the computed standard deviation of each variable and its ratio to the real GDP standard deviation, for the indicated sample periods.

Table 2

CROSS-CORRELATION ANALYSIS WITH REAL GDP

	RGDP (t+k)								
lag k	-8	-6	-4	-2	0	+2	+4	+6	+8
	Sample period: 1980:1 - 2002:4								
RGDP	-0.15	-0.29	-0.03	0.62	1.00	0.62	-0.03	-0.29	-0.15
SRAT	-0.17	-0.06	0.38	0.72	0.52	-0.07	-0.48	-0.38	0.01
LRAT	-0.11	-0.19	0.07	0.51	0.61	0.23	-0.21	-0.34	-0.22
SYDC	0.11	-0.13	-0.44	-0.43	-0.06	0.34	0.43	0.15	-0.25
RYGD	-0.08	-0.29	-0.12	0.41	0.70	0.39	-0.15	-0.34	-0.19
RYGE	0.20	-0.17	-0.30	0.07	0.46	0.31	-0.13	-0.27	-0.10
CQSP	0.20	0.46	0.29	-0.22	-0.58	-0.44	-0.02	0.19	-0.00
	Sample period: 1994:1 - 2002:4								
RGDP	-0.22	-0.42	-0.14	0.54	1.00	0.54	-0.14	-0.42	-0.22
SRAT	-0.18	-0.32	0.12	0.71	0.70	-0.08	-0.63	-0.53	0.00
LRAT	0.27	0.05	-0.12	0.16	0.61	0.29	-0.27	-0.61	-0.52
SYDC	0.44	0.38	-0.25	-0.62	-0.19	0.36	0.43	0.01	-0.49
RYGD	0.22	-0.07	-0.18	0.24	0.77	0.45	-0.22	-0.64	-0.54
RYGE	0.14	-0.19	-0.33	0.11	0.76	0.69	0.08	-0.47	-0.54
CQSP	0.16	0.55	0.37	-0.35	-0.84	-0.44	0.19	0.40	0.04
	Sample period: 1999:1 - 2002:4								
RGDP	-0.45	-0.57	-0.20	0.50	1.00	0.50	-0.20	-0.57	-0.45
SRAT	-0.30	-0.34	0.13	0.74	0.76	0.01	-0.64	-0.72	-0.30
LRAT	0.16	0.01	-0.25	-0.05	0.59	0.36	-0.12	-0.49	-0.47
SYDC	0.40	0.36	-0.30	-0.83	-0.46	0.22	0.63	0.50	0.04
RYGD	-0.01	-0.13	-0.20	0.16	0.75	0.37	-0.18	-0.55	-0.50
RYGE	-0.15	-0.31	-0.30	0.16	0.84	0.53	-0.02	-0.46	-0.50
CQSP	0.31	0.63	0.37	-0.42	-0.95	-0.56	0.12	0.54	0.45

NOTE: all the variables (real GDP, HCPI, interest rates and financial spreads) are in levels. RGDP is real GDP; HCPI is the Harmonized Consumer Price Index; SRAT is the short-term interest rate; LRAT is the long-term interest rate; SYDC is the slope of the yield curve; RYGD and RYGE are the two measures of the reverse yield gap, based, respectively, on dividend yields and earning yields; CQSP is the credit quality spread. The columns report the cross-correlations between the filtered series with a Band-Pass Filter (6,40,8) for the indicated sample periods.

Table 3

CROSS-CORRELATION ANALYSIS WITH CONSUMER PRICES

	HCPI (t+k)								
lag k	-8	-6	-4	-2	0	+2	+4	+6	+8
	Sample period: 1980:1 - 2002:4								
HCPI	0.33	0.48	0.67	0.88	1.00	0.88	0.67	0.48	0.33
SRAT	-0.33	-0.31	-0.14	0.13	0.33	0.37	0.29	0.22	0.23
LRAT	-0.23	-0.25	-0.14	0.07	0.23	0.35	0.40	0.39	0.29
SYDC	0.20	0.16	0.05	-0.10	-0.18	-0.11	0.05	0.12	0.01
RYGD	-0.17	-0.24	-0.18	-0.00	0.15	0.22	0.26	0.29	0.23
RYGE	0.05	-0.03	-0.05	0.03	0.06	0.07	0.13	0.24	0.25
CQSP	0.04	0.22	0.40	0.42	0.27	0.15	0.07	0.04	0.04
	Sample period: 1994:1 - 2002:4								
HCPI	-0.18	0.05	0.43	0.81	1.00	0.81	0.43	0.05	-0.18
SRAT	-0.37	-0.44	-0.29	0.02	0.27	0.33	0.09	-0.16	-0.15
LRAT	-0.25	-0.24	-0.19	-0.04	0.18	0.52	0.62	0.44	0.18
SYDC	0.18	0.25	0.13	-0.06	-0.12	0.13	0.49	0.58	0.33
RYGD	-0.33	-0.40	-0.37	-0.15	0.16	0.50	0.56	0.35	0.15
RYGE	-0.24	-0.48	-0.57	-0.36	0.02	0.34	0.40	0.20	0.03
CQSP	0.11	0.30	0.37	0.30	0.15	0.00	0.04	0.11	-0.05
	Sample period: 1999:1 - 2002:4								
HCPI	-0.42	-0.30	0.14	0.68	1.00	0.68	0.14	-0.30	-0.42
SRAT	-0.22	-0.56	-0.51	0.08	0.73	0.78	0.39	-0.06	-0.27
LRAT	0.02	0.03	-0.21	-0.28	0.25	0.55	0.50	0.11	-0.18
SYDC	0.25	0.61	0.41	-0.27	-0.66	-0.52	-0.12	0.14	0.19
RYGD	-0.06	-0.19	-0.40	-0.33	0.27	0.63	0.54	0.13	-0.16
RYGE	-0.00	-0.28	-0.59	-0.55	0.06	0.52	0.56	0.23	-0.04
CQSP	-0.09	0.46	0.84	0.58	-0.12	-0.54	-0.50	-0.21	0.01

NOTE: all the variables (real GDP, HCPI, interest rates and financial spreads) are in levels. RGDP is real GDP; HCPI is the Harmonized Consumer Price Index; SRAT is the short-term interest rate; LRAT is the long-term interest rate; SYDC is the slope of the yield curve; RYGD and RYGE are the two measures of the reverse yield gap, based, respectively, on dividend yields and earning yields; CQSP is the credit quality spread. The columns report the cross-correlations between the filtered series with a Band-Pass Filter (6,40,8) for the indicated sample periods.

Table 4

REAL GDP FORECASTS: THEIL-U STATISTICS

Sample period: 1994:1-2002:4								
	Forecasting horizon							
Model	1	2	3	4	5	6	7	8
NAIVE	0.79	1.44	2.09	2.75	3.43	4.16	4.89	5.63
ARIMA	0.650	0.542	0.478	0.439	0.409	0.380	0.346	0.299
TV-AR	0.678	0.587	0.535	0.500	0.478	0.456	0.420	0.340
BVAR-M1	0.662	0.560	0.491	0.436	0.401	0.367	0.325	0.234
BVAR-M2	0.656	0.552	0.483	0.430	0.394	0.360	0.317	0.228
BVAR-M3	0.651	0.546	0.478	0.425	0.392	0.362	0.321	0.233
Benchmark	0.624	0.505	0.424	0.364	0.329	0.312	0.291	0.233

NOTE: For the NAIVE model we report the Root Mean Square Error (in percentage points). NAIVE is the random walk model which assumes no variation; ARIMA is an ARIMA (1,1,0) for real GDP and an ARIMA (3,1,0) for inflation. TV-ARs are univariate AR (5) models with time-varying coefficents. BVAR-M1, BVAR-M2, BVAR-M3 are domestic BVAR models including real GDP, HCPI, short-term interest rate and, respectively, the indicated measure of the money supply. The Benchmark model is the same model as BVAR-M2 plus US real GDP.

Table 5

CONSUMER PRICE FORECASTS: THEIL-U STATISTICS

Sample period: 1994:1-2002:4								
	Forecasting horizon							
Model	1	2	3	4	5	6	7	8
NAIVE	0.56	1.01	1.44	1.89	2.35	2.80	3.27	3.70
ARIMA	0.562	0.436	0.338	0.314	0.332	0.357	0.383	0.357
TV-AR	0.602	0.549	0.504	0.512	0.540	0.568	0.586	0.609
BVAR-M1	0.530	0.435	0.305	0.267	0.303	0.324	0.299	0.294
BVAR-M2	0.530	0.427	0.298	0.251	0.287	0.305	0.274	0.266
BVAR-M3	0.534	0.428	0.294	0.252	0.286	0.307	0.274	0.263
Benchmark	0.530	0.427	0.299	0.252	0.288	0.307	0.275	0.268

NOTE: For the NAIVE model we report the Root Mean Square Error (in percentage points). NAIVE is the random walk model which assumes no variation; ARIMA is an ARIMA (1,1,0) for real GDP and an ARIMA (3,1,0) for inflation. TV-ARs are univariate AR (5) models with time-varying coefficents. BVAR-M1, BVAR-M2, BVAR-M3 are domestic BVAR models including real GDP, HCPI, short-term interest rate and, respectively, the indicated measure of the money supply. The Benchmark model is the same model as BVAR-M2 plus US real GDP.

Table 6

ESTIMATED HYPERPARAMETERS

Model	λ_0	λ_1	λ_2	λ_3	λ_4	λ_5	λ_6
TV-AR (rgdp)	1.00	1.0e-7	0.05	1.00	-	0.01	-
TV-AR (hcpi)	1.00	3.5e-8	0.05	1.80	-	0.01	-
BVAR-M1	1.00	4.4e-8	0.15	1.30	0.5	0.01	-
BVAR-M2	1.00	4.4e-8	0.15	1.30	0.5	0.01	-
BVAR-M3	1.00	4.4e-8	0.15	1.30	0.5	0.01	-
Benchmark	1.00	4.4e-8	0.15	1.30	0.5	0.01	0.5

NOTE: λ_0 controls the the decay toward the mean in the law of motion of the coefficients; λ_1 the time variations in the law of motion of the coefficients; λ_2 the general tightness in the prior variance; λ_3 the harmonic decay of the lags in the prior variance; λ_4 the relative tightness of the endogenous variables other the dependent one in the prior variance; λ_5 the prior variance of the deterministic variables; λ_6 the prior variance on the stochastic exogenous variables.

Table 7

BENCHMARK MODEL: RESIDUAL ANALYSIS

	SKEW	KURT	J-B	SKEW	KURT	J-B
Equation	1981:2-1993:3			1981:2-2002:3		
RGDP	-0.30	-0.07	0.77	0.23	0.83	3.27
	(0.40)	(0.92)	(0.68)	(0.39)	(0.13)	(0.20)
HCPI	-0.08	0.15	0.10	-0.09	0.08	0.13
	(0.82)	(0.84)	(0.95)	(0.75)	(0.88)	(0.94)
M2	-0.10	-0.33	0.31	0.29	1.26	7.02
	(0.78)	(0.66)	(0.86)	(0.27)	(0.02)	(0.02)
SRAT	0.18	0.41	0.61	0.34	0.65	3.18
	(0.62)	(0.58)	(0.74)	(0.21)	(0.23)	(0.20)
USGDP	-0.76	1.23	8.17	-0.74	1.50	16.1
	(0.03)	(0.09)	(0.02)	(0.01)	(0.01)	(0.00)

NOTE: The skewness (SKEW) and kurtosis (KURT) statistics include a test of the null hypothesis that each is zero. J-B is the Jarque-Bera test for normality based upon the skewness and kurtosis measures combined. RGDP is euro area real GDP; HCPI is the Harmonized Consumer Price Index; M2 is the money aggregate M2; SRAT is the short-term interest rate; USGDP is the US real GDP. The table reports the computed statistics with their associated p-values in brackets.

Table 8

REAL GDP FORECASTS: THEIL-U STATISTICS

Step	1	2	3	4	5	6	7	8
Model	Sample period: 1994:1 - 2002:4							
NAIVE	0.75	1.37	1.99	2.60	3.22	3.86	4.50	5.17
Benchmark	0.624	0.505	0.424	0.364	0.329	0.312	0.291	0.233
SYDC	0.631	0.516	0.441	0.387	0.358	0.343	0.326	0.281
RYGD	0.623	0.506	0.426	0.367	0.334	0.317	0.297	0.242
RYGE	0.622	0.504	0.423	0.363	0.329	0.313	0.393	0.238
CQSP	0.624	0.506	0.427	0.370	0.334	0.316	0.294	0.236
	Sample period: 1999:1 - 2002:4							
NAIVE	0.88	1.57	2.26	2.94	3.61	4.31	5.02	5.79
Benchmark	0.681	0.558	0.457	0.389	0.357	0.361	0.359	0.261
SYDC	0.687	0.566	0.464	0.390	0.355	0.354	0.356	0.271
RYGD	0.680	0.559	0.457	0.386	0.357	0.361	0.360	0.265
RYGE	0.678	0.557	0.455	0.385	0.356	0.360	0.361	0.266
CQSP	0.678	0.556	0.456	0.390	0.359	0.362	0.358	0.261

NOTE: For the NAIVE model is reported the Root Mean Square Error (in percentage points). NAIVE is the random walk model which assumes no variation. The Benchmark is a BVAR model including real GDP, the Harmonized Consumer Price Index (HCPI), the short-term interest rate, the money M2 and the US real GDP. All the other models include the same variables as the Benchmark plus the indicated financial spread. Specifically, SYDC is the slope of the yield curve; RYGD and RYGE are the two measures of the reverse yield gap, based, respectively, on dividends and earnings; CQSP is the credit quality spread.

Table 9

CONSUMER PRICE FORECASTS: THEIL-U STATISTICS

Step	1	2	3	4	5	6	7	8
Model	Sample period: 1994:1 - 2002:4							
NAIVE	0.59	1.08	1.57	2.05	2.54	3.02	3.49	3.93
Benchmark	0.530	0.427	0.299	0.252	0.288	0.307	0.275	0.268
SYDC	0.532	0.430	0.299	0.251	0.287	0.306	0.274	0.267
RYGD	0.530	0.427	0.299	0.252	0.288	0.307	0.275	0.268
RYGE	0.530	0.428	0.299	0.253	0.289	0.308	0.276	0.269
CQSP	0.529	0.427	0.299	0.252	0.288	0.307	0.275	0.268
	Sample period: 1999:1 - 2002:4							
NAIVE	0.63	1.17	1.72	2.29	2.88	3.48	4.11	4.69
Benchmark	0.587	0.459	0.291	0.236	0.276	0.294	0.224	0.203
SYDC	0.588	0.460	0.288	0.231	0.271	0.290	0.218	0.199
RYGD	0.587	0.459	0.289	0.234	0.273	0.293	0.223	0.203
RYGE	0.588	0.461	0.292	0.237	0.276	0.294	0.225	0.205
CQSP	0.586	0.458	0.290	0.234	0.274	0.292	0.222	0.201

NOTE: For the NAIVE model is reported the Root Mean Square Error (in percentage points). NAIVE is the random walk model which assumes no variation. The Benchmark is a BVAR model including real GDP, the Harmonized Consumer Price Index (HCPI), the short-term interest rate, the money M2 and the US real GDP. All the other models include the same variables as the Benchmark plus the indicated financial spread. Specifically, SYDC is the slope of the yield curve; RYGD and RYGE are the two measures of the reverse yield gap, based, respectively, on dividends and earnings; CQSP is the credit quality spread.

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