Liquidity and Risk Premia of Long-Term Government Bonds - Evidence from EMU and Accession Countries

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Abstract

This PhD thesis contributes to the literature on liquidity and risk premia of long-term government bonds in euro area and in accession countries.

The first paper analyses the effects of shocks to the supply of government bonds in eurozone countries on short-term real interest rates. The interest rate is the rate of return on bonds rather than the rate of return on money. This might be one of the reasons why up to the present the empirical evidence on the liquidity effect of monetary policy shocks has been quite mixed. Therefore, we are looking at the liquidity effect caused by shocks in the government bond supply and the implications these shocks have for fluctuations in short-term interest rates in five euro area countries. We find that in some countries, satisfying a substantial part of their financing need by short-term borrowing, the ex-post real returns investors require on three-month treasury bills are bid up by shocks to the bond supply.

The second paper investigates whether the monetary transmission mechanism of the single eurozone countries has become more homogeneous after 1999. Since the introduction of the euro, a broad strand of literature concerning the homogenisation of the monetary policy transmission process across euro area countries has unfold. This literature has largely ignored the contribution of the collapse of the country risk premia during the build-up to monetary union. We show that risk premia-adjusted yield spreads between ten-year government bonds and three-month treasury bills, which are based on the expectations hypothesis of the term structure, display roughly the same reactions to monetary policy shocks

before and during EMU in euro area countries. This is in contrast to actual yield spreads which clearly change their behaviour with respect to monetary policy shocks between the two sub-periods. The monetary policy transmission mechanism has most likely become more homogeneous since the introduction of the single currency but this has probably mainly been due to the collapse in the country risk premia.

Finally, the third paper examines the relationship between the yield curve and future real economic activity in the Central and Eastern European (CEE) accession countries. We investigate the predictive content of the yield spread for future output growth in the CEE economies looking at two different measures of the yield spread. Namely, the yield spread of five-year government bonds over three-month treasury bills and the yield spread of twelve-month interbank rates over one-month interbank rates. We decompose both measures of the yield spread in an expectations-related and a term premium component and find that innovations to both of the two components indicate future economic expansions in a model including the government spread, whereas they indicate future economic contractions in a model including the interbank spread. This seems to be true on the single country level as well as on the basis of a panel estimation approach.

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Contents

Evidence on the Liquidity Effect in the Euro Area Government Bond Market - A Bayesian VAR Analysis

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The interest rate is the rate of return on bonds rather than the rate of return on money. This might be one of the reasons why up to the present the empirical evidence on the liquidity effect of monetary policy shocks has been quite mixed. Therefore, we are looking at the liquidity effect caused by shocks in the government bond supply and the implications these shocks have for fluctuations in short-term interest rates in five euro area countries. We find that in some countries, satisfying a substantial part of their financing need by short-term borrowing, the ex-post real returns investors require on three-month treasury bills are bid up by shocks to the bond supply.

Keywords : bond supply, liquidity effect
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1 Introduction

We analyse the liquidity effect caused by shocks in the government bond supply and the implications these shocks have for fluctuations in short-term interest rates. This is a deviation from and an extension to the work on the traditional liquidity effect stemming from monetary injections.

The liquidity effect plays a central role in conventional monetary theory and policy. Furthermore, it is an important feature of empirical and theoretical work on the monetary transmission mechanism. In the presence of the liquidity effect, the initial impact of an unanticipated shock to the money supply is to lower nominal and, due to wage and price rigidities, real interest rates for a short period of time. Over a longer horizon agents will adjust their inflation expectations and interest rates will rise again. The theoretical proposition of standard economic textbooks that expansion in monetary aggregates lowers short-term interest rates seems to be widely accepted, yet to date the empirical evidence for it is mixed.

In the $IS-LM$ framework the immediate effect of a monetary expansion is to reduce real rates. If prices are not completely flexible the expansion shifts the LM curve to the right along the downward sloping IS curve resulting in higher output and a lower real rate. When the decline in the real rate more than offsets the effect of the increase in inflation, which probably reacts only with a lag, the nominal rate decreases, too - this is the liquidity effect. In contrast, in a world of flexible-price dynamic general equilibrium models with total price and wage flexibility there would be no role for the liquidity effect because prices would adjust instantaneously during monetary expansions.

In the real world of sticky wages and prices, however, the liquidity effect does affect the way in which monetary policy transmission works. Within the framework of open market policy, fine-tuning operations of central banks provide liquidity to stimulate economic activity. Shocks to the government bond supply can have implications for market liquidity and, thus, for monetary transmission, too. Market liquidity could be increased by both, increasing the money supply or reducing the treasury bond supply. The two seem to be simply two different sides of the same coin.

However, a decrease in the government bond supply, hereafter the total amount of general government securities outstanding, is not equal to an increase in the money supply. This is because the amount of government bonds outstanding can vary apart from the central banks' open market operations due to supply shocks caused by bond auctions. In the auctions in the US and in most euro area countries foreign monetary authorities can roll-over their bond holdings but their roll-over plans are not public information. This causes uncertainty about the actual amount of bonds issued and this uncertainty could manifest itself in investors demanding higher yields when the bond supply is shocked. Even if the quantity of securities allotted were identical to the quantity of securities scheduled for auctioning, the fact that most debt management offices or the respective ministries of finance publicly announce auction schedules only approximately one week prior to auctions causes uncertainty about the amount to be auctioned at the beginning of each specific month during which auctions are to take place. Discrepancies between the agents' forecasts about amounts to be auctioned and actually allotted amounts represent shocks to the total amount of securities outstanding - the bond supply - which could temporarily drive up the borrowing cost for a specific issuer.

The fact that the exact amounts to be auctioned are not known with certainty prior to auctions makes the primary dealers subject to bond-supply risk. The primary dealers that participate in bond auctions commit funds to the bond market before knowing the exact supply and, thus, the exact prices and yields of the bonds. This risk is then passed on to institutional and retail investors on the secondary market. A higher than expected supply in an auction causes prices to fall and yields to rise. This manifests itself in investors in the secondary market requiring higher yields when the bond supply has been subject to shocks.

Given the mixed empirical support for the traditional liquidity effect of monetary policy, our contribution is to analyse the liquidity effect on short-term interest rates deriving from euro area government bond markets. The aim is to analyse the implications of supply shocks of government bonds for market liquidity and for short-term interest rates. As already has been pointed out, the liquidity effect of the bond supply is the reaction of the nominal or the real short-term interest rate to a shock in the outstanding amounts of government securities in any specific country. In our analysis we explain movements in real rates rather than movements in nominal rates because bond investors think in real terms about their investments and try to incorporate inflationary expectations when evaluating their investment opportunities. Public debt management agencies need to calculate issuing costs in real terms, too, when placing the sovereign debt with the public.

We apply a semi-structural vector-autoregressive (VAR) model and a Bayesian VAR (BVAR) in order to study how interest rates respond to monetary policy shocks and to supply shocks in government bond markets in selected countries of the Economic and Monetary Union (EMU) at monthly frequency. The semi-structural VAR leaves the relationships among macroeconomic variables in the system unrestricted but imposes some straightforward contemporaneous identification restrictions on the variables. We build on the work of Jovanovic and Rousseau [15] who find that short-term interest rates react stronger to changes in the supply of marketable treasury securities than they react to changes in the money supply in the United States.

Apart from quantifying the traditional liquidity effect whose causal relationship goes from monetary aggregates to short-term interest rates, we identify and quantify the liquidity effect in the bond market that has been analysed by Grossman and Weiss [13], Jovanovic and Rousseau [15] and Rotemberg [25]. The interest rate being the return on bonds rather than the return on money, the aim is to show that the analysis of the liquidity effect in the bond market is an important building stone in understanding short-run fluctuations in short-term interest rates. We find that in Italy, and to a certain degree in Belgium and in Spain, too, shocks to the bond supply lead to increments in ex-post real returns of three-month treasury bills in the 1980s and the 1990s. This, however, does not seem to be the case in Germany and in France.

Section 2 reviews the main pieces of empirical evidence on the liquidity effect and on monetary policy and summarises their most important results, section 3 explains which kind of data we use and section 4 reviews Jovanovic and Rousseau's [15] methodology of measuring surprises to the growth of the bond supply and estimates the effects on the ex-post real return of treasury bills. Descriptive statistics on the time-series behaviour and interdependencies of the bond supply, the monetary aggregate M1 and interest rates in some euro area countries and in the broader international context in the 1980s and the 1990s are provided in section 5 and the traditional liquidity effect of money supply shocks is analysed in section 6. In order to analyse the dynamics of bond supply shocks, monetary aggregates and the real return on treasury bills in some euro area countries a VAR framework will be introduced in section 7. In section 8 a Bayesian VAR will be employed. The Bayesian VAR overcomes the over-parameterisation problem of the classical VAR by assigning probability distributions to the coefficients. Finally, the conclusions will be drawn in section 9.

2 Review of Literature on the Liquidity Effect

Empirical studies traditionally employed a wide range of specifications, data frequencies and measures for market liquidity. Early studies like Cagan [6] were favourable to the existence of the liquidity effect. In the 1990s and in the new millennium the liquidity effect received support from VAR-based studies such as the ones of Bernanke and Mihov [3], Pagan and Robertson [22], Strongin [29], Leeper, Sims and Zha [17], Christiano et al. [8], Mojon and Peersman [21] and Peersman and Smets [24]. All of these works identified a more or less strong and persistent liquidity effect. Hamilton [14] measured a significant liquidity effect due to variations in the cash balance of the US treasury with the Federal Reserve System on the federal funds rate during and at the end of the two-week reserve maintenance periods.

Leeper and Gordon [16], Christiano [7] and Pagan and Robertson [23], however, raised doubts on the existence of the liquidity effect and reported a vanishing liquidity effect. According to Bernanke and Mihov [3] this might be due to the bias associated with using non-borrowed reserves as a policy indicator. The results of Leeper and Gordon [16] and Christiano [7] stand in the tradition of the rational expectations literature in the 1980s represented, for example, by Mishkin [20] who found that unanticipated changes in the money stock, typically, had no effects on interest rates.

Within the domain of economic theory, Grossman and Weiss [13], Rotemberg [25] and Lucas [19] worked out the liquidity effect of Federal Reserve open-market operations. In these models, in which liquidity-constrained agents require cash or liquidity in order to trade in securities, open-market operations give rise to liquidity effects which generate interest rate behaviour very different from the one predicted by models based on Fisherian fundamentals. In contrast to these models, the sceptical view about the liquidity effect has led to the development of monetized real business cycle models in which persistent money growth leads to an increase in the nominal interest rate.[1]

Leeper et al. [17] analyse models incorporating monetary targeting and interest rate targeting and find that, in line with other monetary policy literature, only a modest portion of output and price level variation in the USA since 1960 can be attributed to shifts in monetary policy. In a four-variable recursive VAR in the CPI, real GDP, the Federal Funds rate and $M1$ they encounter the liquidity puzzle. When $M1$ is shocked the Federal Funds rate decreases initially

[1] Christiano [7] p. 56.

but then bounces back and gains positive territory quickly. Instead, standard textbook macroeconomics would suggest lower interest rates rather than higher interest rates when a monetary aggregate is shocked. On the other hand, when the Federal Funds rate is shocked $M1$ shows the right reaction. Money goes down quickly and persistently. Besides, interest rate tightening leads to an output contraction after six months but creates a price puzzle in the sense that prices increase persistently as a reaction to interest rate tightening whereas one would instead expect them to go down.

The price puzzle could be overcome by including stock prices, long interest rates, exchange rates or commodity prices, which reflect the state of the economy and contain information on the future evolvement of prices in the model. Leeper et al. [17], however, introduce another identification scheme and go away from the recursive identification in order to deal with the price puzzle. If one assumes that it is time-consuming and costly to collect data on prices and on output the money authorities can not react to prices and output immediately. Therefore, $M1$, which is now treated as policy variable will neither be contemporaneously affected by the CPI nor by the output Y. Neither do the CPI or output react to shocks in the interest rate or in money within the period. This is due to the planning processes of the private sector in changing prices and output that take time and are costly. The Federal Funds rate, which is not a policy variable, but treated as informational variable by Leeper et al. [17] responds quickly to all kind of financial market innovations and responds contemporaneously to all other variables. The fact that the interest rate can react immediately to the CPI and to output, but the central bank is not conceded this capacity to react immediately leaves room for discussions about the central bank procedures and about its capabilities of reaction.

The model with the second identification scheme of Leeper et al. [17], incorporating lags to allow for the decision-making process of private agents and the central bank, delivers straightforward impulse responses to shocks in $M1$. A policy tightening in the form of lower $M1$ leads to lower prices, an initial but only slight decrease of output which then returns to its previous level and an increase in the interest rate, which then returns to its initial level after about one year. A striking feature is that $M1$ decreases very strongly and persistently while output and the price level hardly react. What is problematic in this model are, however, the positive reaction of $M1$ and the very strong positive reaction of the CPI to interest rate shocks.

The model imposing constraints on private agents' reactions to the interest rate and $M1$ combined with the central bank's reactions to output and prices generates a liquidity effect of $M1$ on the interest rate for about one year before the nominal interest rate returns to its initial level. Instead, the above model with a recursive identification scheme displays the wrong reaction of the interest rate to $M1$ shocks.

Clearly, both models have drawbacks as displayed by their impulse responses and, what is more relevant for our study, they are not very helpful in resolving the liquidity puzzle. The recursive model delivers the right reaction of $M1$ to the Federal Funds rate shocks but the wrong reaction of the Federal Funds rate to $M1$ shocks. $M1$ and the Federal Funds rate in the non-recursive model react vice versa. This means that the two models neither can explain whether positive $M1$ shocks decrease interest rates, nor whether positive interest rate shocks decrease $M1$.

One problem in identifying the liquidity effect might be that $M1$ is not a very good proxy for measuring the policy stance of the Federal Reserve because it is demand determined up to a certain degree. Private sector decisions on prices and on output have direct effects on $M1$ and, therefore, $M1$ can not be controlled in the short run by the Federal Reserve. This would suggest the use of aggregates such as borrowed reserves, non-borrowed reserves or total reserves, which are more directly controlled by the central bank, as a measure for liquidity.

Christiano et al. [8] follow up this approach. They distinguish between three different benchmark identification schemes, each of which is recursive. The three benchmark identification schemes are one model in which the federal funds rate is the policy variable, one in which non-borrowed reserves (NBR) are the policy variable and a third one, with which we will not deal here, in which both NBR and total reserves are policy variables.

Variables are determined in a block recursive way within the quarter. First goods market variables are determined, then policy variables and lastly money market variables. Recursiveness requires that a set of variables is predetermined relative to the policy variable. When deciding on monetary policy the central bank looks only at predetermined variables and monetary policy shocks can be identified using the fitted residuals of ordinary least squares regressions of the policy variable on the predetermined variables. This is in contrast to many papers in the literature that drop the assumption that the central bank looks only at predetermined variables and which require further identifying assumptions.

The model variables included are: a monetary aggregate, output, the price level, a commodity price index, the federal funds rate, NBR and total reserves.

The block-recursiveness of the benchmark models implies that the output, the prices and the commodity price index are not shocked by the policy variables within the quarter. Their sample data refers to the US economy from 1965 to 1995.

In all three models Christiano et. al. [8] assume that the central bank contemporaneously reacts to output, the price level and a commodity price index when it takes its decisions on the policy variable. This assumption, according to them, is at least as realistic as the assumption, taken by many other researchers, that monetary authorities can't react to the price level and to output within the period. Their data frequency being quarterly, this might be an innocuous assumption. Even if the central bank has no precise within quarter information on output and prices at its disposal it can check weekly and monthly statistics on many important and meaningful variables such as unemployment figures, retail sales and wages. The assumption that the central bank takes into account prices and output within the period might be a bigger issue when monthly data are used. The higher the frequency of the data, the more difficult it is for the monetary authorities to get the development of prices and output right.

In Christiano et. al.'s [8] Federal Funds rate model a shock to the Fed Funds rate leads to an increase in the Fed Funds rate, a fall in output, a fall in the price level after a lag of six quarters, a fall in the commodity price index with a later rebound of the commodity prices after six quarters, a negative response of NBR which, however, returns to zero after only three quarters, a slight fall in TR and, finally, a persistent fall in the monetary aggregate $M1$. There is a strong liquidity effect at work in this model. A contractionary shock to the Fed Funds rate generates negative responses of $M1$ and of NBR.

It should be very informative to analyse the NBR model as NBR innovations are supposed to reflect mainly exogenous shocks to monetary policy, whereas innovations to broader monetary aggregates are heavily influenced by private sector shocks to money demand.

In Christiano et. al.'s [8] NBR model a contraction in NBR leads to a fall in output which is less pronounced that the fall in output caused by a rise in the interest rate in the federal funds model. The negative shock to NBR leads to a fall in the price level, a fall in the commodity price index similar to the one in the federal funds rate model, a rise in the federal funds rate, a strong fall in total reserves, and a larger fall in $M1$ than in the first model. So, the liquidity effect seems to be even stronger in the NBR model.

Both models, the Federal Funds rate and the NBR model display clear evidence of the traditional liquidity effect by which positive shocks to the Federal Funds rate and negative shocks to NBR generate negative responses of $M1$.

For the euro area Peersman and Smets [24] analyse the liquidity effect. Peersman and Smets [24] analyse the monetary transmission mechanism for the euro area as a whole on the basis of synthetic euro area data from 1980 till 1998 at a quarterly frequency. They apply an identified VAR framework to study the macroeconomic effects of monetary policy shocks in the euro area. They find that their macroeconomic effects are similar to the ones found in US data and that they are surprisingly stable over time.

Using various standard identification schemes their main findings are that temporary rises in nominal and real interest rates lead to temporary falls in output and to a sluggish response of prices which tend to fall only several periods after the monetary policy shock takes place. They identify an immediate but subdued liquidity effect of a temporary rise in the short-term interest rate on $M1$. $M3$, containing as well interest-bearing assets decreases, however, more gradually. Share prices fall, house prices are sluggish and long-term interest rates rise slightly in response to a short rate hike. Furthermore, they find evidence of an appreciating exchange rate.

Peersman and Smets [24] analyse a VAR containing a vector of exogenous and a vector of endogenous variables. The vector of exogenous variables consists of the world commodity price index, US real GDP and the US short-term nominal interest rate. These variables, on which the euro area variables have no feedback, should control for developments in the world economy and deal with the prize puzzle. The vector of endogenous variables includes euro area real GDP, euro are consumer prices, the monetary aggregate $M3$, the nominal short-term euro area interest rate and the real effective euro exchange rate.

Peersman and Smets [24] use a Cholesky-decomposition of the exogenous and the endogenous variables vector. So, monetary policy shocks have no contemporaneous effects on output, prices and money but on the exchange rate. The exchange rate is last in the Cholesky ordering as in a large economy like the euro area the exchange rate is assumed to have no contemporaneous effect on monetary policy.

The results of Peersman and Smets [24] VARs are in line with other work in the area. A monetary contraction lowers output after two quarters and leads to an appreciation of the exchange rate. Prices react sluggish initially but decrease after a couple of periods.

For a model for the US economy their impulse responses are very similar to the euro area ones with the only exception that policy shocks in the US on average have a bigger magnitude than in the euro area and a bigger effect on output and prices.

Regarding the liquidity effect, Peersman and Smets [24] find an initial negative reaction of $M1$ to monetary policy shocks. However, after four quarters $M1$ becomes positive before returning to the baseline after nine quarters. The reaction of $M3$ is initially less pronounced but over time shows more persistence. Initially, monetary tightening causes substitution of funds that bear no interest with time deposits or money market funds, which are included in the broad monetary aggregate $M3$. Therefore, $M3$ as a whole hardly shows a reaction in the first six quarters. Then, after seven quarters $M3$ declines persistently.

The reaction of $M1$ casts some doubt on the liquidity effect in the synthetic euro area model. This is in contrast to Christiano et al.'s [8] findings of a negative reaction of $M1$ to Federal Funds rate shocks for the US. However, this clear contrast in the results could very well be due to aggregation problems in the synthetic euro area data. Countries are assumed to have central banks with identical monetary policy reaction functions. Clearly, this is not a very satisfying working assumption and could generate incorrect results on the existence of the liquidity effect.

Therefore, Mojon and Peersman [21] follow a different line of reasoning than Peersman and Smets [24]. Instead of looking at an area wide model they account for the fact that members of the euro area are heterogeneous when it comes to the monetary transmission mechanism. When the effects of monetary policy on economic activity and on the price level differ from country to country it is essential to estimate central bank reaction functions at the country level. Mojon and Peersman [21] analyse the transmission mechanism of monetary policy in ten countries that are members of the euro area. They classify countries into three groups according to the degree by which their monetary policy was constrained by the EMS. Germany as the EMS anchor constitutes a group of its own, while the second group consists of the Netherlands, the monetary union of Belgium and Luxembourg, and Austria which had de facto a fixed exchange rate towards the German Mark and no autonomy in monetary policy. Finally, there is the third group of countries headed by France and Italy that maintained a certain degree of autonomy in monetary policy.

Mojon and Peersman's [21] results are qualitatively similar to Peersman and Smets [24]. A contractionary monetary policy shock leads to a temporary fall in output, to a gradual decrease in the price level and to an initial decrease of

$M1$ for most countries. Although EMS countries were subject to asymmetric monetary policy shocks, these country level results resemble the results of the aggregate euro area economy of Peersman and Smets [24].

Mojon and Peersman [21] introduce different identification schemes for each country group. For Germany they choose one of the benchmark models from Peersman and Smets [24]. The model consists of a vector of exogenous variables including the world commodity price index, US real GDP and the US short-term nominal interest rate. The choice of these variables should solve the price puzzle and control for world demand and world inflation. The exogenous variables have contemporaneous effects on the vector of endogenous variables which is identical to the one as in Peersman and Smets [24].

For the group consisting of Austria, Belgium and the Netherlands, German output, German prices, the German short-term interest rate and the bilateral exchange rate with Germany are included into the vector of endogenous variables because economic developments in these countries are dominated by developments in neighbouring Germany and, therefore, autonomous monetary policy shocks are very unlikely. Hence, the monetary policy shock is identified as the shock to the German interest rate in these countries.

The respective endogenous macro variables of the three countries have no contemporaneous effects on the German variables. The macro variables included are GDP, the CPI, a short-term interest rate, the real effective exchange rate and $M1$.

Constituting an interest rate anchor for the EMS, the short-term German interest rate is included in the endogenous variables vector for the third country group. The omission of the German interest rate could cause a price puzzle because a domestic interest rate increase on the grounds of an interest rate increase in Germany could be wrongly associated with exchange rate depreciation. This would put upward pressure on prices in contrast to the expected decrease of the CPI in contractionary periods.

Mojon and Peersman [21] measure domestic monetary policy shocks as the deviation of domestic short-term interest rates from the German rate and identification is achieved by a standard recursive identification scheme.

For all countries a contractionary domestic monetary policy shock leads to a fall in output and in prices. Interestingly, these effects are much less pronounced for the euro area as a whole than on the single country level. Less consistence is found in the effect of a policy shock on the exchange rate. The exchange rates of Belgium and the Netherlands do not react to monetary policy shocks. This reflects the credibility of these countries' EMS parities. The exchange rates of all other countries with the exception of Italy and Spain do, however, appreciate in response to a tightening of policy. Finally, Italy and Spain experience an exchange rate puzzle in the sense that their currencies depreciate when policy is tightened. This might be due to a market sentiment postulating that the monetary tightening represents in reality a manoeuvre to protect a depreciating currency.

For countries that try to fix their bilateral DM exchange rates Mojon and Peersman [21] find just a limited role for monetary aggregates in the model. Therefore, their evidence on the existence of the liquidity effect is quite mixed. $M1$ hardly reacts to monetary policy shocks in France, Greece, Ireland and the euro area as a whole. For Spain, Finland and Italy $M1$ initially declines but returns to its old level after approximately ten quarters. For Germany, Austria, Belgium and the Netherlands $M1$ initially declines but then rebounds and increases above its starting level after five quarters. As could have been expected, $M3 - M1$ which is mainly driven by time deposits and money market funds increases in all countries with the exception of Greece.

On a whole, the recent empirical results are quite favourable to the existence of a liquidity effect of monetary expansions. However, they show that the evidence that has been collected on the liquidity effect depends crucially on the identification schemes and on the sample periods which have been employed. The fact that there is some doubt on the existence of the liquidity effect led Jovanovic and Rousseau [15] to analyse the liquidity effect caused by shocks to the bond supply. The liquidity effect of the bond supply is the reaction of the nominal and real short-term rates to shocks to the bond supply.

3 The Data

We analyse the effect of bond supply shocks on real short rate fluctuations in the five eurozone countries with the biggest amount of government securities outstanding. Namely, these are in alphabetical order Belgium, Germany, Italy, France and Spain. The yields of three-month treasury bills and the amounts of government securities outstanding were retrieved from the national central banks' websites. data on consumer prices, monetary aggregates and industrial production were taken from DataStream.

RR is the monthly real ex-post return of an agent that has bought a treasury bill with three-month remaining maturity on the primary market and holds the bill until maturity. The real return is constructed as the difference between the average monthly gross allotment rate for three-month maturities and the three-month inflation rate measured by the seasonally-adjusted consumer price index. Money $M1$ is the logarithm of the real value of the seasonally adjusted monetary aggregate $M1$ retrieved from DataStream. The bond supply BS is the log of the real marketable central government debt outstanding. IP is the log of seasonally adjusted real industrial production, which we use instead of real GDP since it is readily available as a monthly series. Finally, P is the logarithm of the seasonally adjusted consumer price index. The IMF commodity price index and the German one year bond yield will be treated as exogenous variables in the VAR. The world commodity price index contains important information on the future path of inflation and the German interest rate contains important information on the future course of the European Monetary System's (EMS) countries' interest rates.

We use the total amount of government securities outstanding instead of the narrower measure of the bond supply for our analysis. The bond supply per se would be the amount of government securities outstanding minus the amount held by the respective country's central bank. The estimation results are, however, robust to the use of the bond supply per se instead of total government securities. We illustrate this for the case of Italy in appendix A. Therefore, we will refer to the two terms as if they were identical. By the same token, we interchange the expressions government securities and treasury securities. Strictly speaking, one would need to distinguish between these two stocks because general government debt might as well include the debt of state railways, the post offices, the social security system or local or state governments additionally to the treasury debt. Interchanging general government securities and central government securities should not pose a problem in our choice of countries, since in all five countries central governments are implicitly expected to bail-out local and state governments in times of financial turmoil.

4 Earlier Evidence on the Liquidity Effect in the Bond Market

Jovanovic and Rousseau [15] quantify the liquidity effect on interest rates stemming from supply-risk in the bond market. At the monthly frequency, surprise bond purchases by the Federal Reserve System raise bond prices and reduce yields. Furthermore, the residual bond supply-risk due to random rollover plans of foreign financial and monetary institutions prevents the market makers from precisely predicting the actual bond supply. The supply-risk adds between 10 and 40 basis points to the standard deviation of the real interest rate on three-month treasury bills. Towards the end of the Clinton administration it was thought that in the US the bond supply-risk might increase in the future as the gradual paying down federal debt would have implied that it would have become harder to expand treasury bill issues to accommodate unexpectedly large rollover demands from foreign financial and monetary institutions. This problem does not pertain to EMU with its higher and more persistent levels of government debt.

In contrast to monetarist theory, money does not seem to have had clear effects on real interest rates in the US since the 1970s. Jovanovic and Rousseau [15] suggest that, the interest rate being the return on bonds and being pinned down by the supply of bonds rather than by the supply of money, the bond supply is in any case the better measure in order to analyse the liquidity effect than is money.

The importance of analysing the variance in the bond supply shows up in the fact that since 1980 growth in NBR has reduced short-term interest rates, but not as strongly as a contraction of treasury securities would have done. Surprisingly, over the whole period from 1960 to 1999 growth in NBR is positively correlated with short-term rates. As would have been expected, growth in marketable treasury securities is positively correlated with short-term rates for both the 1961 to 1999 and the 1980 to 1999 sub-period.

The evidence from Jovanovic and Rousseau [15] seems to suggest that at least for the US the liquidity effect in the bond market is a more straightforward vehicle to explain changes in short-term rates than the traditional liquidity effect via a monetary aggregate. Their conclusions do not change when they look instead at surprises in the bond supply and in NBR.

The amount of marketable treasury securities outstanding and money in the form of non-borrowed reserves could be expected to be negatively correlated. When the central bank buys treasury securities by the means of open-market operations $NBRs$ increase. Jovanovic and Rousseau [15] , however, find only for the period between 1980 and 1999 a negative correlation between the growth in real per capita supply of outstanding treasury securities and NBR. For the period from 1961 to 1999 as a whole they find a positive correlation.

The main result of [15] consists in quantifying surprises to the bond supply and to inflation and measuring their impact on movements of the real ex-post return of three-month treasury bills. They measure surprises as the one-step-ahead forecast errors of a system of forecasting equations which comprise the monthly real return to a US three-month treasury bill, the monthly US inflation rate based on the CPI for all urban consumers, the monthly growth rate of the real level of marketable US government securities outstanding and, finally, the monthly real return on the S&P500 index. The forecasting equations are based on a rolling VAR with a 36-month estimation window to allow for the inclusion of contemporary economic developments. Eventually, the surprises to the real return, to the bond supply and to inflation are pooled over the rolling sample of the VAR and, finally, the surprises to the real return are regressed on the surprises of the other two variables.

Treasury bills being zero-coupon bonds, the nominal return $R_{t,t+1}$ needed to calculate the real ex-post return $r_{t,t+1}$ for the forecasting equations is based on a zero-coupon bond with price P_t at date t:

$$R_{t,t+1} = \frac{1 - P_t}{P_t} \tag{1}$$

The real ex-post return on this bond is

$$r_{t,t+1} = (\frac{1}{P_t}(\frac{1}{1 + \pi_{t,t+1}}) - 1), \tag{2}$$

where $\pi_{t,t+1}$ is the monthly inflation rate between dates t and $t+1$. Taking logs yields

$$\ln(1 + r_{t,t+1}) = -\ln P_t - \ln(1 + \pi_{t,t+1}) \tag{3}$$

For any small number ε, $\ln(1+\varepsilon) \approx \varepsilon$. Thus,

$$r_{t,t+1} \approx R_{t,t+1} - \pi_{t,t+1} \tag{4}$$

The nominal return $R_{t,t+1}$ on a zero-coupon bond equals the nominal market interest rate $i_{t,t+1}$. Then, if the real return $r_{t,t+1}$ can be broken up in an expected component and in a surprise component we can approximate,

$$r^u_{t,t+1} = i^u_{t,t+1} - \pi^u_{t,t+1} \tag{5}$$

This means that the surprise to the ex-post real return $r^u_{t,t+1}$ is made up of the surprise component to the nominal return $i^u_{t,t+1}$ and the surprise component to inflation $\pi^u_{t,t+1}$.

Concerning the information set of the agents, at the beginning of period t agents do know the realised inflation $\pi_{t-1,t}$ between dates t and $t-1$, the realised real T-bill return $r_{t-1,t}$ between dates t and $t-1$ and the real return from the S&P 500 between dates t and $t-1$. However, agents have no exact information on the growth rate $g_{t-1,t}$ of the bond supply between dates t and $t-1$. This information arrives too late to be included in their information set. Therefore, agents have no precise forecasts about the price P_t of a bond. When agents commit funds to the bond market before the date t auctions take place they do not know yet the growth of the bond supply during period $t-1$.

[15] assume a liquidity effect of the bond supply on the price of bonds of the form

$$i^u_{t,t+1} = \alpha g^u_{t-1,t}, \tag{6}$$

where $g^u_{t-1,t}$ is the surprise component to the bond supply growth between dates t and $t-1$. 6 into 5 yields

$$r^u_{t,t+1} = \alpha g^u_{t-1,t} - \pi^u_{t,t+1} \tag{7}$$

The occurrence of bond supply risk implies that agents do not know the date t supply of bonds at the time at which they form their expectations of the new auction price P_t and, thus, do not know the future nominal interest rate $i_{t,t+1}$. The reason that agents at time t can not precisely know the bond supply and, thus, the auction price is due to the residual supply-risk inherent to US treasury auctions. Apart from the published auction amount of bonds in an auction the actual supply depends on the rollover plans of foreign financial and monetary institutions which submit non-competitive bids. Their rollover plans not being known to the public, results in the creation of residual supply-risk.

The problem then boils down to estimating the regression

$$r^u_{t,t+1} = \alpha + \beta g^u_{t-1,t} - \gamma \pi^u_{t,t+1} + \epsilon_t, \tag{8}$$

where $r^u_{t,t+1}$, $g^u_{t-1,t}$ and $\pi^u_{t,t+1}$ are the surprises to the monthly ex-post real return of a treasury bill with a remaining maturity of three months, to the growth in the real amount of marketable treasury securities outstanding and to the monthly inflation rate. These surprises are measured by one-step ahead forecast errors from three forecasting equations which include the real return, the growth of treasury securities, inflation, the real return to the S&P 500 index and, finally, a time trend. Eventually, the surprises to the real return, to the growth rate of treasuries and to inflation are pooled over the rolling sample of a VAR with monthly observations from January 1920 through December 1999 and then the surprise to the real return is regressed on the other two surprises as shown in equation 9. T-statistics are in parenthesis.

$$\begin{aligned} r^u_{t,t+1} &= \underset{(0.60)}{0.0001} + \underset{(7.45)}{0.0274}\, g^u_{t-1,t} - \underset{(-17.31)}{1.082}\, \pi^u_{t,t+1} \end{aligned} \tag{9}$$

The surprise to the supply of government securities $g^u_{t-1,t}$ has a significant positive impact on the ex-post real return $r^u_{t,t+1}$ of three-month T-bills. Replicating the rolling VAR and the forecasting equations of Jovanovic and Rousseau [15] for our dataset from January 1966 through March 2000 we get slightly different results for the regression of the surprises as can be seen in equation 10:[2]

$$\begin{aligned} r^u_{t,t+1} &= \underset{(0.33)}{0.0011} + \underset{(0.29)}{0.0027}\, g^u_{t-1,t} - \underset{(-18.41)}{0.2432}\, \pi^u_{t,t+1} \end{aligned} \tag{10}$$

5 The Evidence on the Variability of the Bond Supply

Figure 1 illustrates the correlations of monthly $M1$ growth and monthly bond supply growth with the real return on three-month treasury bills in the upcoming three-months and the correlations of $M1$ growth and bond supply growth with three-month nominal treasury bill yields in the upcoming three-months for the sample from 1980 to 2002.

The real and nominal returns of the upcoming three months are chosen for the reason that afterwards the situation of primary dealers and investors who buy three-month treasury bills and hold them until maturity will be analysed. The fact that primary dealers commit funds to bond auctions before they know

[2] In contrast to Jovanovic and Rousseau (2001), we do not assume that dividends will be reinvested when calculating the real return to the S&P 500.

the exact amount that will be auctioned and, thus, before they know the price and the yield of the bonds to be auctioned makes them subject to the bond-supply risk which they then pass on to the investors on the secondary market. Shocks to the monthly bond supply might imply lower bond prices and primary dealers and investors might react in demanding higher yields when buying bonds at the beginning of each three-month holding period.

Therefore, on the one hand, one would associate growth in the bond supply with higher short-term nominal and real interest rates while, one the other hand, one would expect growth in the money supply to be negatively correlated with short-term real and nominal interest rates. Looking at table 1 containing the correlations, Germany, France and Spain display the expected negative correlation between the growth rate of money $M1$ and the nominal short-term rate, while Belgium and Italy display positive correlations. The real short-term rate, instead, is negatively correlated with monetary growth in all countries with the only exception being Italy. Concerning the bond supply growth, it is positively correlated with the real short-term rate only in Belgium and Italy and it is positively correlated with the nominal rate in Belgium, Italy and France. In contrast, the bond supply growth in Germany displays a clearly negative correlation with both, real and nominal short-term rates.

The liquidity effect seems to be at work in Germany, Spain and France where we find a negative correlation between $M1$ and the short-term nominal interest rate but not in Belgium and in Italy. Belgium and Italy could be an interesting case to study because they are the only two countries which display the wrong - positive - sign for the correlation between monetary growth and nominal short-term interest rates on the one hand, paired with the presumably correct, i.e. positive, correlation between bond supply growth and real and nominal short-term rates on the other hand.

The positive and, therefore, intuitively wrong correlation between the nominal short-term interest rate and money $M1$ for Belgium and Italy could imply that these countries are special in the sense that in these countries variations in the bond supply, displaying at least the correct sign in contrast to the variations of money supply, might have more explanative power for the fluctuations in nominal interest rates than variations in the money supply have. On the other hand, the positive correlation between $M1$ growth and the nominal short-term rates could as well be due to private sector shocks that have nothing to do with monetary policy shocks. Private sector shocks could be at the heart of increases to money demand and faster growth in broad monetary aggregates

Table 1: Correlations of treasury securities' and M1's monthly growth rates with three-month real and nominal returns

	BE	DE	ES	FR	IT
$\mathrm{corr}(r_{t,t+1}, g_t^{BS})$	0.04	−0.14	−0.03	−0.11	0.06
$\mathrm{corr}(r_{t,t+1}, g_t^{M1})$	−0.11	−0.18	−0.23	−0.05	0.02
$\mathrm{corr}(i_{t,t+1}, g_t^{BS})$	0.10	−0.27	−0.00	0.08	0.05
$\mathrm{corr}(i_{t,t+1}, g_t^{M1})$	0.05	−0.14	−0.08	−0.15	0.15

like $M1$ and, thus, drive up short-term interest rates. Therefore, one needs to disentangle monetary policy shocks from private sector monetary shocks.

The same caveat is obviously in place regarding the positive correlations between the bond supply growth and the three-month nominal and real return in Belgium and Italy. The positive correlations could be due to private sector shocks which have nothing to do with structural shocks to the bond supply. Therefore, one needs to identify the structural shocks in the economy in order to be capable of making assertions on the causal relationship between bond supply risk and the ex-post real return on three-month treasury bills. Doing exactly this in a structural BVAR for Italy, we find evidence that shocks to the bond supply raise the real return of three-month treasury bills. For France and Germany, however, shocks to the bond supply seem to decrease the real return. Only Belgium and Spain follow Italy in the presence of the liquidity effect in the bond market up to a certain degree. For Belgium and Spain we find evidence for a slight increase of the real returns in response to shocks of the bond supply.

Of course shocks to the bond supply can only be relevant in explaining interest rate fluctuations if the growth rate of the amount of government securities outstanding is not simply constant but has a certain degree of variability. Here, it is interesting to compare the variability in the bond supply to the variability in the money supply as money is usually considered the variable to have the most immediate effect on short-term interest rates.

Figures 1 and 2 illustrate the monthly volatility of the real $M1$ and the real treasury securities growth rates exemplary for Italy and for Germany[3]. It can be seen that especially in Germany the volatility of the growth rates of treasury securities has a magnitude in the order of the volatility of monetary growth rates. Despite of this, the volatility in the bond supply is a phenomenon that has been rarely analysed as to date.

[3]We apply the Hodrick-Prescott filter on these two series.

Figure 1: Twelve-month rolling standard deviations of monthly growth rates of real M1 and real treasury securities in Italy

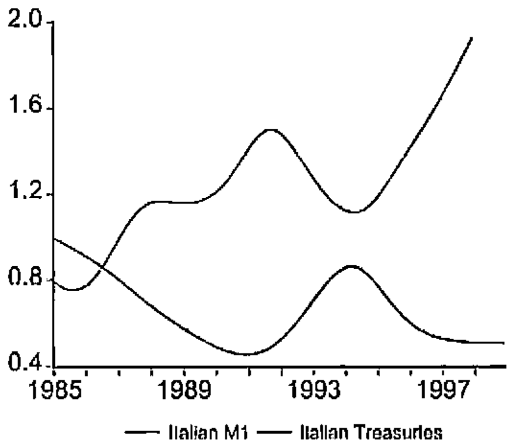


Figure 2: Twelve-month rolling standard deviations of monthly growth rates of real M1 and real treasury securities in Germany

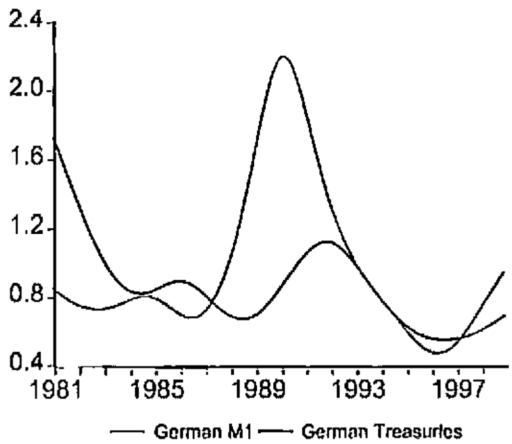


Table 2: Monthly growth rates of real M1 and real treasury securities

	BE	DE	ES	FR	IT
μ^{M1}	0.31	0.43	0.28	0.09	0.69
μ^{BS}	0.59	0.75	1.52	1.03	0.44
σ^{M1}	1.54	1.28	0.86	0.78	1.41
σ^{BS}	5.06	1.14	5.20	7.33	0.91

Table 2 displays the means and standard deviations of monthly growth rates of real $M1$ and real treasury securities for all five countries of our study from 1980 to 2002. For most countries in our sample the mean and the standard deviation of the monthly growth rates of treasury securities have been exceeding the ones of $M1$.

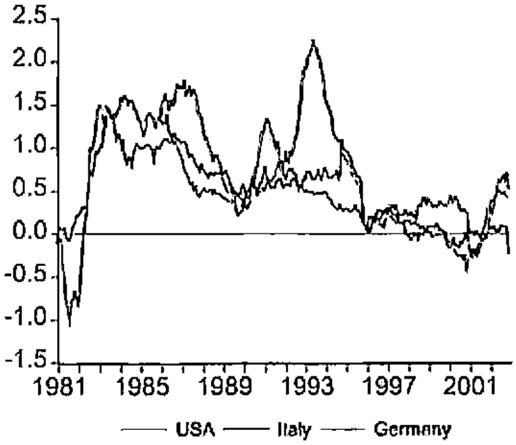
Figure 3 depicts how the monthly growth rates of real treasury securities compare in the US, in Italy and in Germany. The graphs imply that in all three countries there is considerable variability in the amount of treasuries outstanding. This variability could put significant upward pressure on the real returns of the treasury securities. In fact [15] find that in the US shocks to the bond supply cause positive shocks on the real returns on three-month treasury bills and in this paper we will illustrate that in Italy shocks to the total amount of treasuries outstanding lead to slight increases in the real return on three-month treasury bills. This might be an exclusively Italian phenomenon because the variability in the growth rate of treasury bills outstanding might historically have been higher in Italy than it has been in other euro area countries and the share of treasury financing needs covered by issuing treasury bills might have been higher in Italy than for example in France or in Germany.

For Belgium and for Spain we find similar results like for Italy. Bond supply shocks seem to exert an upward pressure on the ex-post real returns of three-month treasury bills in these countries, too.

6 Money Supply Shocks

What does the evidence on the traditional liquidity effect regarding shocks to monetary aggregates look like for the five euro area countries Belgium, Germany, France, Italy and Spain? We look at a four-variable structural VAR from Leeper et. al. [17] discussed in section 2 in order to compare the dynamics in the data of the five euro area countries to a standard benchmark model of monetary

Figure 3: 12 month moving average of the monthly growth rates of treasury securities outstanding in the US, in Italy and in Germany



policy analysis. Regarding the identification of the VAR, we impose one further homogeneous restriction, i.e. $\alpha_{21} = 0$ which implies that price shocks have no contemporaneous effect on output. This restriction is motivated by the existence of a certain lag in private agents' decisions in reaction to CPI shocks.

For the general motivation of this identification scheme please refer to the section 2 where we give a detailed explanation of the different identification schemes of Leeper et. al. [17]. The lagged reactions of the monetary authorities to private sector variables and the sluggish reactions of private agents to money and interest rates give rise to the following contemporaneous restrictions on the coefficients of the variables:

$$\mathrm{A}e_t = \mathrm{B}u_t, \tag{11}$$

$$\text{where } \mathrm{A}e_t = \begin{pmatrix} 1 & \alpha_{12} & 0 & 0 \\ 0 & 1 & 0 & 0 \\ \alpha_{31} & \alpha_{32} & 1 & \alpha_{34} \\ 0 & 0 & \alpha_{43} & 1 \end{pmatrix} \begin{pmatrix} e_t^{CPI} \\ e_t^{IP} \\ e_t^{TB3m} \\ e_t^{M1} \end{pmatrix},$$

$$\text{and } \mathrm{B}u_t = \begin{pmatrix} \beta_{11} & 0 & 0 & 0 \\ 0 & \beta_{22} & 0 & 0 \\ 0 & 0 & \beta_{33} & 0 \\ 0 & 0 & 0 & \beta_{44} \end{pmatrix} \begin{pmatrix} u_t^{CPI} \\ u_t^{IP} \\ u_t^{TB3m} \\ u_t^{M1} \end{pmatrix}$$

Although our main objective is to analyse the effect of shocks to the supply of treasury securities on the ex-post real return of three-month treasury bills we begin to estimate the four-variable structural model from Leeper et al. [17] for the data of our five eurozone economies to get evidence on the traditional liquidity effect of monetary innovations on short-term interest rates. To this end, we use the three-month treasury bill yield $TB3m$ instead of the real return and we leave aside the amount of treasuries outstanding for the time being.

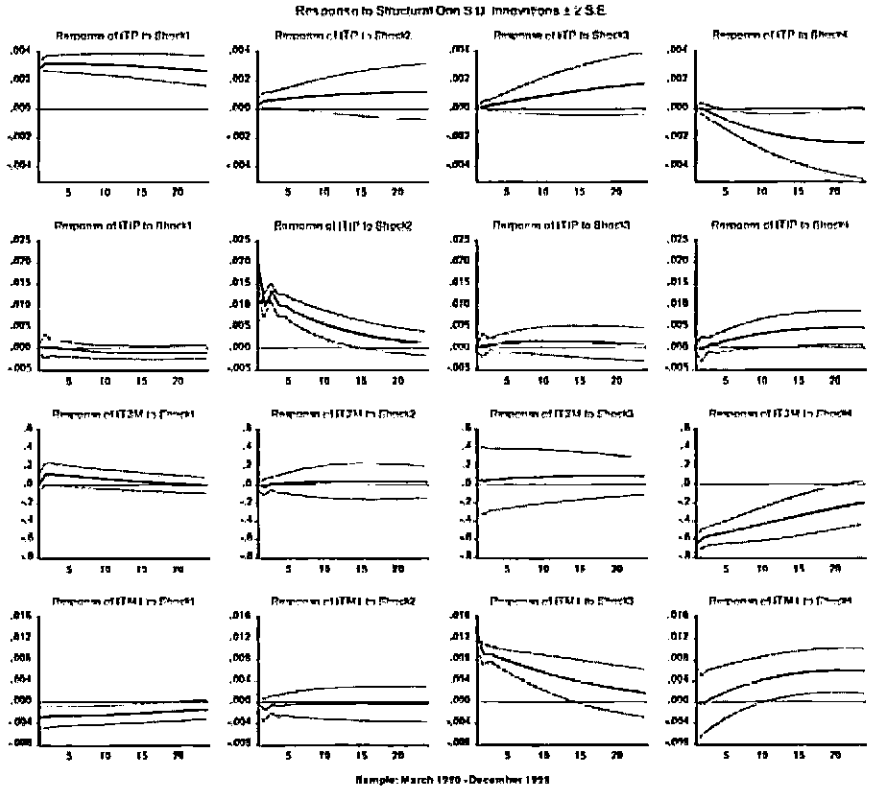
Figure 4 shows the impulse responses for this model for the case of Italy. Consumer prices react positively to the output shock and show the expected negative reaction to the interest rate shock. In all four other countries prices show a negative reaction to interest rate shocks, too. For Spain, however, the reaction is quite weak.

Italian output initially does not react much neither to the interest rate nor to $M1$. However, it shows a slightly positive reaction to money after several periods. This is very similar to the reactions in Belgium, Germany and France but in contrast to Spain, where output reacts negatively to $M1$ shocks.

The interest rate, instead, displays a strong liquidity effect in Italy when $M1$ is shocked but it hardly reacts to the other variables. $M1$ shows a positive reaction to interest rate shocks but the magnitude of the reaction is not very strong.

The liquidity effect for the other countries, for which we do not print the impulse responses, can be read off the tables 3 and 4 displaying structural one standard deviation innovations to annualised three-month treasury bill rates and $M1$ with their respective standard errors in parenthesis. For France and Belgium the three-month rate decreases when $M1$ is shocked. In Germany and Spain instead, the three-month return seems to respond positively to monetary shocks. This could happen via an expectations effect when markets expect higher monetary growth to feed through to higher inflation and interest rates rise in anticipation of higher policy rates. A second explanation could be that in Germany and in Spain private demand shocks are responsible for the positive effects of money shocks on the interest rate. Private demand shocks could raise

Figure 4: Four-variable structural VAR for Italy



money demand and, therefore, increase interest rates. In France and Belgium the traditional liquidity effect seems to be at work like in Italy. Here, money supply shocks are the dominant force and lower the short-term nominal interest rate.

Regarding, instead, the shock of the interest rate on $M1$ we find that the shock lowers $M1$, as it should do, in all countries. The reaction is strongest in Belgium, Italy and in Germany.

Table 3: Impulse responses of annualised three-month interest rates to one standard deviation innovations in M1

Period	BE	DE	ES	FR	IT
1	-0.342818	0.238199	0.343984	-0.585103	-0.670219
	(0.03034)	(0.01251)	(0.02433)	(0.03049)	(0.03613)
6	-0.379783	0.286907	0.293130	-0.504091	-0.515617
	(0.04801)	(0.03475)	(0.05052)	(0.06583)	(0.06673)
12	-0.267166	0.191344	0.233406	-0.360976	-0.404968
	(0.06019)	(0.04175)	(0.06450)	(0.09194)	(0.10020)
18	-0.172610	0.111064	0.169842	-0.245911	-0.296320
	(0.07187)	(0.04452)	(0.07306)	(0.10231)	(0.11719)
24	-0.101270	0.047377	0.110291	-0.156092	-0.198957
	(0.07437)	(0.04168)	(0.07779)	(0.10149)	(0.12004)

underlying structural model is system (11)

Table 4: Impulse responses of M1 to one standard deviation innovations in annualised three-month interest rate

Period	BE	DE	ES	FR	IT
1	-0.014053	-0.011820	-0.005362	0.007193	0.012798
	(0.00061)	(0.00050)	(0.00034)	(0.00042)	(0.00060)
6	-0.010443	-0.009704	-0.008672	0.006469	0.007437
	(0.00108)	(0.00139)	(0.00163)	(0.00101)	(0.00112)
12	-0.009092	-0.007692	-0.008319	0.005020	0.004902
	(0.00221)	(0.00198)	(0.00240)	(0.00148)	(0.00184)
18	-0.008073	-0.006296	-0.008088	0.003823	0.003004
	(0.00303)	(0.00232)	(0.00330)	(0.00179)	(0.00219)
24	-0.007317	-0.005376	-0.007848	0.002845	0.001635
	(0.00339)	(0.00239)	(0.00417)	(0.00195)	(0.00227)

underlying structural model is system (11)

7 Bond Supply Shocks and Real Returns

We analyse the effects of bond supply shocks on the real ex-post return of three-month treasury bills of the five euro area countries Belgium, Germany, France, Italy and Spain at a monthly frequency. Even if the amount of securities outstanding might not be the most sophisticated measure of market liquidity from a market micro structure perspective, choosing the five countries with the biggest amount of government securities outstanding in the euro area allows us to abstract at least to a certain degree from micro structure issues, which, by having an impact on yields, could distort our analysis.

Bond supply shocks are caused by the nature of the way in which ministries of finance and debt management offices are auctioning and syndicating government bonds. When bonds are to be auctioned, auction schedules are typically released only one week prior to auctions indicating the quantity to be auctioned. Due to rollover from foreign monetary authorities or, in some countries, due to securities purchased directly by national central banks the scheduled quantity could differ substantially from the actual quantity allotted. Even if the quantity announced and the quantity allotted were identical, agents still would be subject to supply-risk at a monthly frequency because the amounts to be auctioned are only released few days prior to the auctions and agents, therefore, face a certain degree of uncertainty about the amounts to be auctioned at the beginning of any specific month.

Ideally, we would like to measure the shocks to the growth rates of the amount of three-month treasury bills outstanding and analyse their impact on investors' required real returns. This would provide us with a measure of how costly high variance in amounts issued or, more generally, a lack of transparency regarding the emission of treasury bills could be for the public debt financing. An elevated cost in public debt financing could manifest itself in investors demanding higher yields when facing higher market risk on primary and secondary markets in the presence of shocks to the amounts of securities issued and in the presence of auctioning or syndication methods that are characterised by a lack of transparency or reliability.

However, constructing time series on the outstanding amount of specific maturities of treasury bills would be possible only at prohibitive cost. Therefore, we need to content ourselves with measuring shocks to the total amount of government bonds and bills outstanding and the implications they have for real ex-post returns on three-month treasury bills.

7.1 Cholesky Identification

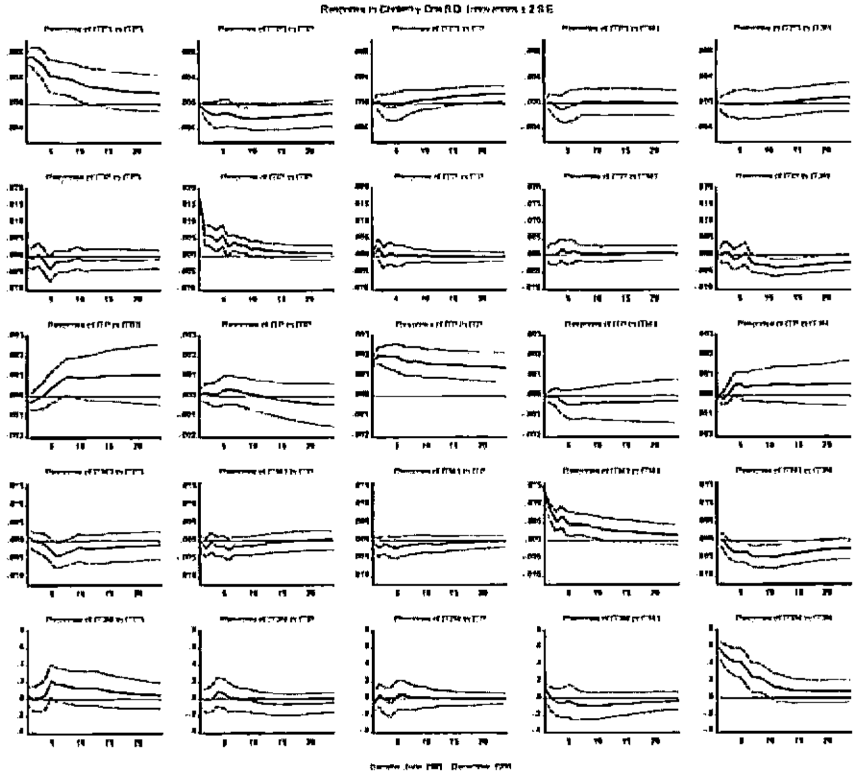
We kick-off our analysis of the effects of bond supply shocks with a simple Cholesky identification scheme estimating a VAR in Italian industrial production, the Italian real bond supply, Italian real $M1$, the monthly yield on an Italian treasury bill with three months remaining maturity and the Italian consumer price index. We include the three-month treasury bill yield rather than the ex-post real return to compare the impulse responses for the Italian data to the benchmark models of Leeper et. al. [17] and we use industrial production rather than output because of its availability as a monthly series. The quarterly series for GDP carry too much noise into the impulse responses when converted into monthly frequencies. The comparison to models like the one of Leeper et al. [17] could already reveal the possible opportunities from introducing the variable bond supply in VARs analysing monetary shocks and interest rate fluctuations.

A Cholesky decomposition of these variables with the Cholesky-ordering BS, IP, P, $M1$, $3m$ yields the impulse responses shown in figure 5. The bond supply is ordered first because usually it does not respond within the period to the other macro variables. The same is true for industrial production and consumer prices which react only with a certain lag to variables like $M1$ and the three-month interest rate. The IMF commodity price index and the German one year nominal interest rate are included as exogenous variables in the system.

The Italian bond supply reacts negatively to industrial production but does not seem to react to shocks in any of the other variables. In some of the other countries, for which we do not display the impulse response functions of the Cholesky decomposition, the bond supply reacts to some of the other variables, too. However, no general pattern can be established of how the bond supply reacts to prices, interest rates or money. Clearly, the most interesting response for us is the reaction to the interest rate. For the case of Italy, the bond supply reacts very sluggishly to interest rate shocks but the reaction seems to go at least in the right direction.

Industrial production shows hardly a reaction to $M1$ Italy. This seems to reject the monetarist view that monetary aggregates are capable of moving production. Neither does industrial production initially respond to the interest rate shock. It displays the expected negative reaction only from the 6th period onwards. Leeper et al. [17] have suggested a negative shock of output, the variable they use instead of industrial production, to shocks in the federal funds rate in a four-variable model including CPI, output, the federal funds rate and

Figure 5: Impulse responses of a Cholesky decomposition including the nominal yield on Italian three-month treasury bills



$M1$. This seems to be a more straightforward result. In the four countries other than Italy, industrial output reacts positively to $M1$ and the magnitudes of the impulse responses are bigger.

The CPI reacts positively to the bond supply after three to four periods and negatively to $M1$. There is a positive reaction of the CPI to the interest rate after two to three periods - the price puzzle. The price puzzle is no exclusively Italian phenomenon. It shows up in the other four countries, too. Consumer prices react positively to interest rate shocks, while they would be supposed to display negative reactions.

Italian $M1$ reacts negatively to industrial production, to the bond supply, to the three-month interest rate and to the price level. The impulse responses to shocks in industrial production and the price level display, however, quite large error bands. The contraction of $M1$ in response to the interest rate is quite strong and persistent. This liquidity effect has the correct sign in contrast to the one that Leeper et al. [17] find for the US and in contrast to the result we get from the four-variable model excluding the bond supply in 4. Although there is an identification issue which has to be borne in mind here it would seem that the inclusion of a new variable - the bond supply - in the monetary model has helped to resolve the liquidity puzzle.

In the other four countries $M1$ reacts positively to shocks in industrial production but the responses to other variables vary from country to country. Of particular interest are the responses of $M1$ to interest rate shocks. $M1$ contracts in all countries even if with a smaller magnitude and with less persistency than for the Italian case.

Finally, the Italian three month interest rate hardly reacts to industrial production but it displays a clearly positive reaction to the bond supply after being initially inert for four months. After five periods a one percentage point bond supply shock raises the three-month rate by some 20 basis points. The interest rate initially reacts positively to $M1$ but after two months it displays the expected negative reaction. The reaction is, however, much mor muted than in the model excluding the bond supply. The interest rate hardly reacts to industrial output and to consumer prices.

The liquidity effect of shocks to the monetary aggregate seems to be rather small for Italy. A one percentage point shock to $M1$ leads to a decline of the three-month rate of roughly 15 basis points only after nine periods. For Italy the short-term interest rate seems to be more sensitive to bond supply innovations than to monetary innovations.

With the exception of Germany, the other countries replicate the negative reaction of the interest rate to monetary shocks although the initial reaction in Belgium is positive like in Italy. Regarding the reaction of the interest rate to bond supply shocks, interest rates react negatively in Germany and France, while in Belgium and Spain they initially react positively but then decline after few periods.

The most striking result is that the Cholesky identification scheme creates a price puzzle. When interest rates are shocked, the CPI goes up. This is in contrast to the four variable model presented in equation 11 above which generates the correct responses of the CPI to interest rate shocks for Belgium, Germany, Spain and France but not for Italy. Another possibility would be that the introduction of the new variable government bond supply causes the prices to react in the wrong way. Nice in the Cholesky scheme is that industrial production goes down in all five countries due to the interest rate shock instead of staying unchanged or even going up as in the four-variable structural VAR before. Another positive feature of the model is that for Italy and France the reaction of $M1$ to the interest rate now comes in with the right sign in contrast to the four-variable structural VAR, so that finally all countries display the correct reaction of $M1$ to interest rate shocks. Furthermore, the identification scheme generates the correct liquidity effect of $M1$ on interest rates in all countries, with the exception of Germany where the liquidity puzzle found in the four-variable structural VAR in equation 11 continues to persist.

7.2 Structural VAR

After introducing the bond supply into the model we proceed to estimate a structural VAR to identify the structural shocks of the bond supply, the monetary aggregate $M1$ and the short-term interest rate to the system. The Cholesky identification scheme presented above was too rigid, in the sense that it restricted the interaction between the three-month interest rate, $M1$ an the bond supply too much without allowing for contemporaneous feedback between these variables, especially between the interest rate and $M1$.

The nominal yield on three-month treasury bills $TB3m$, money $M1$, the government bond supply BS, industrial production IP and the consumer price index P constitute the structural model

$$\mathrm{A}\begin{pmatrix} TB3m_t \\ M1_t \\ BS_t \\ IP_t \\ P_t \end{pmatrix} = \mathrm{C(L)}\begin{pmatrix} TB3m_{t-1} \\ M1_{t-1} \\ BS_{t-1} \\ IP_{t-1} \\ P_{t-1} \end{pmatrix} + \mathrm{B}\begin{pmatrix} u_t^{TB3m} \\ u_t^{M1} \\ u_t^{BS} \\ u_t^{IP} \\ u_t^{P} \end{pmatrix}, \qquad (12)$$

where the matrix A defines the contemporaneous interaction between the variables of the model, and C(L) is a matrix lag polynomial. u_t^{TB3m}, u_t^{M1}, u_t^{BS}, u_t^{IP} and u_t^{P} are the structural disturbances to the model's variables. The matrix B defines which structural shocks will hit which variables in the system.

The reduced form of the structural model (12) can be written as

$$\begin{pmatrix} TB3m_t \\ M1_t \\ BS_t \\ IP_t \\ P_t \end{pmatrix} = \mathrm{A}^{-1}\mathrm{C(L)}\begin{pmatrix} TB3m_{t-1} \\ M1_{t-1} \\ BS_{t-1} \\ IP_{t-1} \\ P_{t-1} \end{pmatrix} + \begin{pmatrix} e_t^{TB3m} \\ e_t^{M1} \\ e_t^{BS} \\ e_t^{IP} \\ e_t^{P} \end{pmatrix}, \qquad (13)$$

where the vector $\mathbf{e}_t' = \begin{bmatrix} e_t^{TB3m} & e_t^{M1} & e_t^{BS} & e_t^{IP} & e_t^{P} \end{bmatrix}$ contains the VAR residuals.

The identification of this structural model is achieved by orthogonalising the structural disturbances, imposing the theory-based restriction that macroeconomic variables do not contemporaneously react to monetary or financial market variables and imposing the restriction that the policy makers and private agents react only with lags to macroeconomic variables. These restrictions are in line with models in the monetary policy literature which study effects of monetary policy on macroeconomic variables with datasets of monthly frequency. The general motivation is that the private sector variables react sluggishly and that the central bank reacts with a lag to macro variables.

Thus, the relationship between the VAR residuals and the structural disturbances can be written as

$$\mathbf{A}\mathbf{e}_t = \mathbf{B}\mathbf{u}_t. \qquad (14)$$

$$\text{where } \mathrm{A}e_t = \begin{pmatrix} 1 & \alpha_{12} & \alpha_{13} & \alpha_{14} & \alpha_{15} \\ \alpha_{21} & 1 & \alpha_{23} & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} e_t^{TB3m} \\ e_t^{M1} \\ e_t^{BS} \\ e_t^{IP} \\ e_t^{P} \end{pmatrix}$$

$$\text{and } \mathrm{B}u_t = \begin{pmatrix} \beta_{11} & 0 & 0 & 0 & 0 \\ 0 & \beta_{22} & 0 & 0 & 0 \\ 0 & 0 & \beta_{33} & 0 & 0 \\ 0 & 0 & 0 & \beta_{44} & 0 \\ 0 & 0 & 0 & 0 & \beta_{55} \end{pmatrix} \begin{pmatrix} u_t^{TB3m} \\ u_t^{M1} \\ u_t^{BS} \\ u_t^{IP} \\ u_t^{P} \end{pmatrix}$$

The first equation in the system posits that the interest rate on three-month treasury bills, being a variable which is updated in real time by the financial markets, is assumed to be contemporaneously affected by all other variables in the model. Money contemporaneously affects the interest rate - this is the liquidity effect. Furthermore, the interest rate on T-bills is affected by bond supply changes. In the special case that bond supply changes constitute mainly changes in long-term bonds rather than changes in T-bills the effect on the short-term interest rate can, however, be expected to be very limited.

As can be seen from the second equation money does react to the interest rate and to the bond supply within the period. Due to lags in central bank decisions and in private sector dispositions it does not react to industrial production and to the CPI within the period.

The third equation in the system implies that the bond supply is contemporaneously only influenced by its own shocks. The interest rate, money, industrial production and the CPI do not contemporaneously affect the bond supply because the quantity of new government bonds issued does not vary due to these variables at high frequencies. The assumption that the debt issuing agency does not react to shocks to the interest rate might be true up to a certain extent because once an auction and the amount of T-bills to be auctioned have been decided upon and announced it is very unlikely that the auction will be cancelled or the amount to be auctioned will be modified on the account of an interest rate shock. At most there might be small adjustments in the amount to be auctioned in order to accommodate developments in the markets.

The fourth equation restricts the contemporaneous reaction of private agents' production decisions to the interest rate, to money, to the bond supply and to the CPI to zero. Due to costly and time-intensive dispositions in the private sector, industrial production does not react contemporaneously to interest rates, money, the bond supply and the CPI.

Finally, the fifth equation reflects the often postulated view in the empirical monetary transmission literature that the aggregate price level responds very little to monetary contractions and expansions. In this equation sluggishness in the private sector motivates the restrictions that the price level does not react to interest rates, money, the bond supply and industrial production.

The commodity price index is added as exogenous variable to the structural VAR as it contains important information on the future course of inflation. Due to its importance in accounting for EMS member countries short-term interest rates, the German one-year interest rate figures as exogenous variable in the system.

7.2.1 Nominal Returns

Figure 6 displays the impulse responses for Italy for this model. The interest rate reacts negatively to shocks to money $M1$ and it reacts positively to bond supply shocks. This is in line with both models studied above - the four-variable SVAR and the Cholesky identification scheme including the bond supply as additional variable.

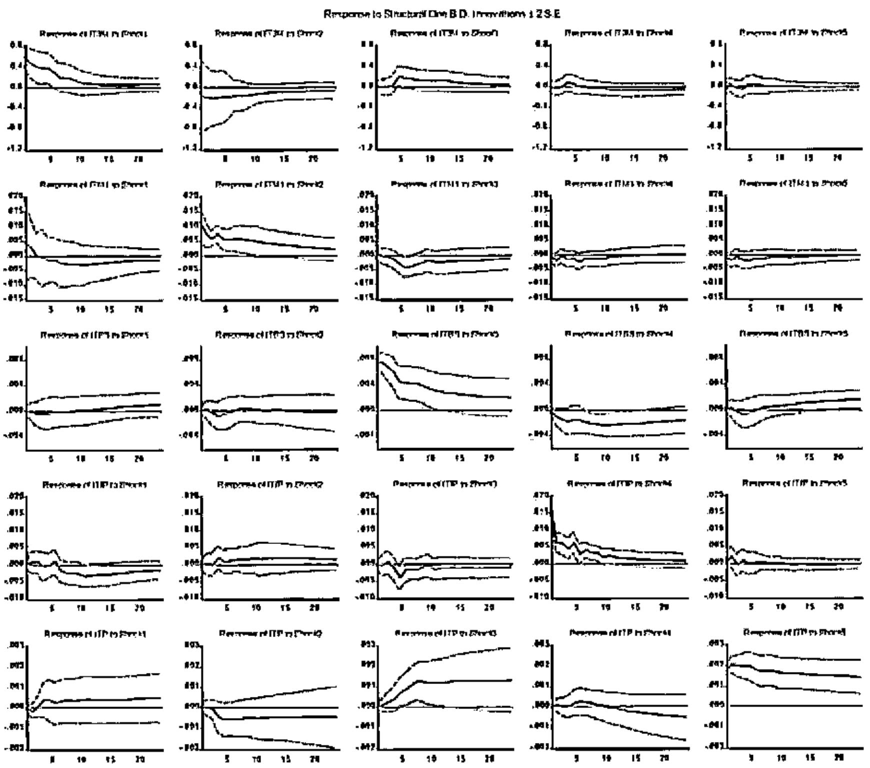
Regarding the other four countries, for which we do not display the impulse responses, we observe the liquidity puzzle, in the sense that the nominal short-term rate reacts positively to money supply shocks, only for France. Although the response of the German three-month rate to money supply shocks becomes positive after about ten periods, too. This again confirms the above results.

$M1$ in Italy displays the correct sign in its reaction to the interest rate shock after some periods. This, however, cannot be said for Belgium, Germany and Spain where money then reacts positively to interest rate shocks. For Belgium and Germany this is in contrast to the first two models which generated the correct response for money to interest rate shocks.

Finally, the Italian consumer price index displays a positive reaction to the interest rate shock. This price puzzle manifests itself in the other four countries, too. The price puzzle had already been observed in the model with the Cholesky identification scheme while it was not present in the four-variable structural model excluding the bond supply.

Regarding the traditional liquidity effect we can state that the three models

Figure 6: Structural VAR for Italy including the nominal annualised yield on Italian three-month treasury bills



including annualised three-month nominal interest rates are generally supportive of the view that shocks to the money supply lower nominal short-term interest rates. The message of the three models at hand is, however, not totally equivocal. Important exceptions are the strong positive reaction of the German and the Spanish interest rates in model 11, the positive reaction of the German rate in the Cholesky identification scheme and the strong and positive reaction of the French interest rate in the structural VAR including the bond supply in model 14.

Concerning the reaction of the short-term interest rate to bond supply shocks, the results of model 14 suggest that in Italy, and to a very limited degree in Belgium and in Spain, the interest rate reacts positively to the bond supply shocks. In Italy the bond supply shock on the interest rate has roughly the same magnitude as the M1 shock has on the interest rate. The Cholesky identification scheme yields qualitatively the same results.

7.2.2 Real returns

Now it is time to introduce the ex-post real return instead of the nominal yield into the model. The difference between the two measures being the change in the consumer price index during the three months in which agents hold a treasury bill with a remaining maturity of three months. The contemporaneous interaction between the real return and the other four variables of the model might give rise to different dynamics and interaction between the model variables than in the nominal yield model above.

The real return for an investor in a buy-and-hold strategy is defined by the yield to maturity at the purchase of a three-month treasury bill minus the inflation in the following three months. An investor buying a new issue in an auction or on the secondary market will, therefore, react to information on the possible future path of inflation. News about the commodity price index, the CPI, industrial production and $M1$ will influence his perception about possible levels for future inflation and, therefore, the yield which he will require from his investment. Information on possible shocks in the bond supply will influence his valuation of the bond he is about to acquire independently from the expected inflation path.

The contemporaneous shocks given to the real return are identical to the ones to the nominal yield of treasury bills. The real return is contemporaneously shocked by the commodity price index, by $M1$, by the bond-supply, the consumer price index and by industrial production. A shock to $M1$ will most likely lead to higher prices of treasury bills and, therefore, lower nominal and real returns. Shocks to the bond supply could push down bond prices on secondary markets or, when they are caused by roll-over decisions, on primary markets and, therefore, push up yields and real returns. These reactions could, however, depend on where on the maturity curve the supply shocks takes place. Supply shocks on the long end may not have as strong effects on short-term rates as supply shocks on the short end would have. Finally, a rise in the CPI implies higher future inflation and should raise nominal yields, but investors that have already taken a long position in the market will see their real returns decrease by inflation surprises.

The identifying restrictions of the real yield model are identical to the ones in system (14). We display the corresponding impulse responses for the case of Italy only in figure 7, while we display the impulse responses of three-month real returns to all shocks for the five countries in tables 5 to 9.

The real return is now expressed as the return accruing during three months in contrast to the annualised nominal return used in the nominal yield models. In Italy, the real return reacts negatively to $M1$ and displays a slightly positive reaction to the bond supply after a lag of eight months. Although not being significant in this model, the lagged reaction of the real return to bond supply shocks reminds of the lag in the impulse response of the nominal yield to the bond supply in figure 6 and the reaction in the Cholesky identification scheme in figure 5.

$M1$ increases as a reaction to interest rate shocks and decreases after CPI shocks. The increase of M1 is reminiscent to the structural VAR in the nominal yield of figure 6 where M1 increased for few periods before gaining negative ground. The bond supply, instead, displays a negative reaction to the shock in industrial production and a positive, but sluggish, reaction to shocks in the CPI.

By and large, the models in the nominal and in the real return deliver the same qualitative results for Italy. Even with the very high inflation at the beginning of the Italian sample in the second half of the 1980s, the impulse responses of the real and the nominal models are very similar. The only notable exception is the reaction of the price level to interest rate shocks. The CPI reacts positively to shocks of the nominal three-month rate - the 'price puzzle' - in figure 6 but it shows the correct, negative, reaction to three-month real returns in figure 7.

Figure 7: Structural VAR for Italy including the three-month ex-post real return on Italian three-month treasury bills

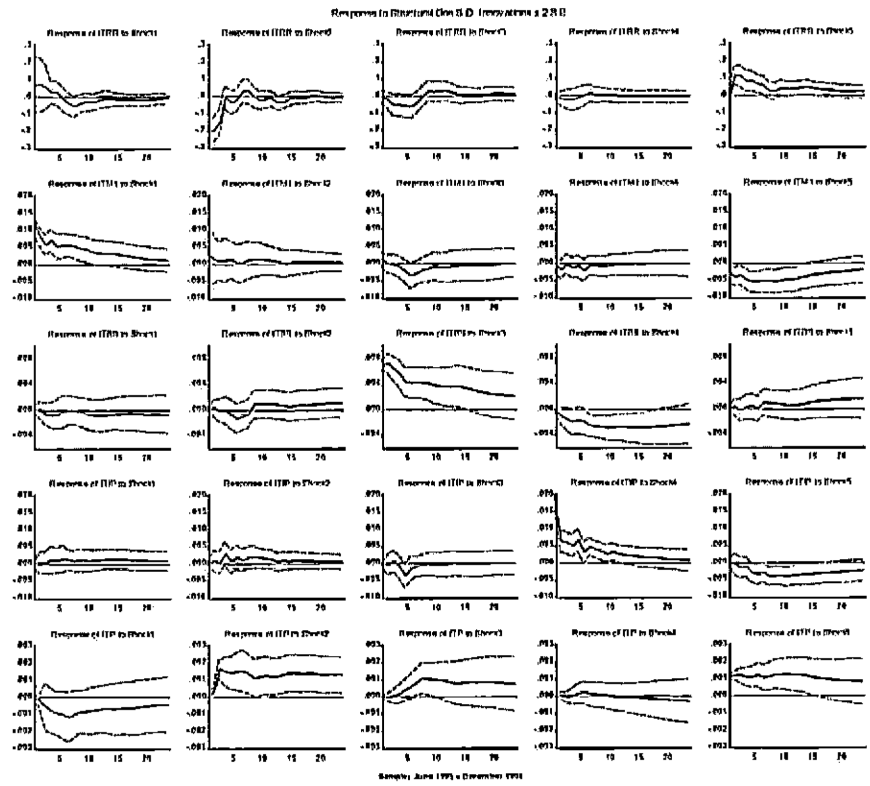


Table 5: Italy - impulse responses of three-month real returns to one standard deviation structural innovations

Period / Shock	RR	M1	BS	IP	P
1	0.070303	-0.203992	-0.002795	0.001134	-0.013692
	(0.08186)	(0.03085)	(0.0169)	(0.00566)	(0.01687)
6	-0.011144	-0.020184	-0.062432	0.000777	0.067486
	(0.03544)	(0.03269)	(0.03357)	(0.03064)	(0.02710)
12	-0.023412	-0.012718	0.031797	-0.004186	0.038460
	(0.02301)	(0.02249)	(0.02886)	(0.01895)	(0.02060)
18	-0.018257	-0.005666	0.003723	-0.005124	0.032196
	(0.01712)	(0.01692)	(0.02077)	(0.01797)	(0.01933)
24	-0.011589	-0.006104	0.009911	-0.007184	0.019546
	(0.01570)	(0.01232)	(0.01970)	(0.01815)	(0.01870)

structural model is represented by equation (14)

Table 6: Belgium - impulse responses of three-month real returns to one standard deviation structural innovations

Period / Shock	RR	M1	BS	IP	P
1	0.297393	-0.216924	0.029750	0.019595	0.009317
	(0.18593)	(0.25513)	(0.02300)	(0.02223)	(0.01944)
6	0.004028	0.008746	-0.030657	-0.041962	-0.031016
	(0.03793)	(0.03788)	(0.03630)	(0.03634)	(0.03016)
12	0.008299	-0.022509	0.009498	-0.012857	0.002348
	(0.02885)	(0.02062)	(0.01289)	(0.01837)	(0.01474)
18	-0.000664	-0.012605	0.008230	-0.013550	0.008390
	(0.01633)	(0.01031)	(0.00774)	(0.01008)	(0.00893)
24	-0.004947	-0.010398	0.005671	-0.011232	0.012269
	(0.01318)	(0.00848)	(0.00718)	(0.00736)	(0.00789)

structural model is represented by equation (14)

Table 7: Germany - impulse responses of three-month real returns to one standard deviation structural innovations

Period / Shock	RR	MI	BS	IP	P
1	0.014956	0.211144	-0.041897	-0.001248	0.001517
	(0.09803)	(0.01149)	(0.01303)	(0.00823)	(0.00998)
6	0.021007	-0.022840	-0.015387	0.062170	0.008366
	(0.02596)	(0.02483)	(0.02015)	(0.02284)	(0.01707)
12	0.018624	-0.005561	-0.012210	0.023788	-0.004273
	(0.01508)	(0.01918)	(0.01166)	(0.01577)	(0.01074)
18	0.008280	-0.006005	-0.008697	0.004417	0.003902
	(0.00795)	(0.00854)	(0.00731)	(0.01122)	(0.00711)
24	0.004440	-0.004977	-0.005970	0.000256	-0.000379
	(0.00576)	(0.00490)	(0.00682)	(0.00927)	(0.00646)

structural model is represented by equation (14)

Table 8: France - impulse responses of three-month real returns to one standard deviation structural innovations

Period / Shock	RR	MI	BS	IP	P
1	0.106206	0.122791	0.061399	0.000000	0.031217
	(0.04731)	(0.04073)	(0.01979)	(0.00000)	(0.01750)
6	-0.002879	-0.061466	-0.013207	0.059284	-0.002119
	(0.04042)	(0.03376)	(0.03300)	(0.03975)	(0.02722)
12	0.021011	-0.010792	-0.004263	-0.030206	-0.020014
	(0.02907)	(0.02839)	(0.03388)	(0.03152)	(0.02655)
18	0.017250	0.009655	0.002209	0.009617	0.022030
	(0.02858)	(0.02180)	(0.03000)	(0.02298)	(0.02143)
24	-0.003439	-0.015388	-0.007971	0.004526	-0.009131
	(0.02869)	(0.02026)	(0.02679)	(0.02040)	(0.01685)

structural model is represented by equation (14)

Table 9: Spain - impulse responses of three-month real returns to one standard deviation structural innovations

Period / Shock	RR	M1	BS	IP	P
1	0.241593	0.183120	-0.005976	-0.002219	0.107066
	(0.04047)	(0.05523)	(0.02745)	(0.02188)	(0.02829)
6	-0.078480	-0.122446	0.026533	0.003547	0.035208
	(0.05721)	(0.06100)	(0.05250)	(0.05017)	(0.03807)
12	0.006799	-0.009296	-0.015370	0.007571	0.008414
	(0.03725)	(0.04974)	(0.02920)	(0.02557)	(0.02401)
18	0.004679	0.017347	-0.002540	0.007341	-0.006184
	(0.01968)	(0.03689)	(0.01774)	(0.01404)	(0.01598)
24	0.003180	0.031350	0.006027	0.008054	-0.007005
	(0.01454)	(0.03120)	(0.01626)	(0.01351)	(0.01112)

structural model is represented by equation (14)

Tables 5 to 9 indicate that shocks to the bond supply exert upwards pressure not only on nominal bond yields, as demonstrated in section 7, but as well on real bond yields. This seems to be the case for Italy and Belgium but probably as well for Spain and France.

8 Bond Supply Shocks and the Bayesian VAR Approach

The number of coefficients in a VAR grows as a quadratic function of the number of variables included. Therefore, even VARs with only five or six variables are heavily parametrised models. Heavily parametrised models in conjunction with scant sample information which is very random, however, cause an overfit and result in models that reflect random noise rather than systematic empirical variability. The application of arbitrary exclusion restrictions founded on economic theory or the use of the Bayesian approach to estimation represent potential solutions to this over-parameterisation problem of a VAR.

A priori one often does not know whether it is feasible to set values for specific coefficients equal to zero. However, there are many situations in which knowledge of model coefficients is not totally absent. In this case the Bayesian approach avoids too many zero restrictions and allows the available information to be expressed more realistically through allocation of probability distributions to the model's different coefficients.

In the Bayesian approach each of the model's coefficients are assigned probability distributions. Mean and variance of these distributions control the likelihood of a given coefficient taking on a particular value. Hence, stochastic prior information on coefficients allows a reasonable range of uncertainty about parameter values.

In order to overcome the overparametrisation problem and to get rid of an excess of random noise in the impulse response functions of our structural VAR we employ a Bayesian VAR (BVAR) to analyse the effects of bond supply shocks on the three-month real returns of treasury bills.

8.1 The Minnesota prior

A Minnesota prior incorporates the fact that a random walk is a good proxy for the behaviour of economic variables over time and the prior takes into account the fact that a more recent path of a variable yields more information about its behaviour than its more distant past. Normal distributions are assigned to each coefficient of a lag so that the mean is one for the coefficient of the first own lag and zero for the other lags. Coefficients on longer lags are more likely to be close to zero than coefficients on shorter lags. The variance will be lower for other lags than for own lags of and will decrease for more distant lags.

However, the specification of a complete normal prior on a VAR would be intractable and a few well-chosen hyperparameters are used instead. The hyperparameters of which we make use are the mean vector of coefficients on the first own lags, the importance of other lags with respect to own lags and the lag decay.

When variable j refers to the jth variable in the VAR and equation i to the equation whose dependent variable is i, the information needed to construct the prior is:

First, the mean of the prior distribution for the first own lag in each equation. We use a mean of zero for the prior on all coefficients except the first own lag in each equation on which a prior mean of one is placed. The reason for this is that many economic time series follow a random walk process. The alternative is a mean of zero on series which can be represented by white noise. In our system we do not have any variables that behave like white noise.

Second, the standard deviation $S(i, j, l)$ of the prior distribution for lag l of variable j in equation i for all i,j and l. The standard deviation is scaled by the standard errors of variable j in equation i to control for the different scale of the variables in the system.

The standard priors restrict the standard deviation function to the form:

$$S(i,j,l) = \{\gamma g(l) f(i,j)\} s_i/s_j, \tag{15}$$

where γ is the overall tightness, which is the standard deviation on the first own lag and $g(l)$ is the tightness on lag l relative to lag 1. The lag tightness expresses the importance of longer lags relative to the first lag.

Third, $f(i,j)$ is the tightness on variable j in equation i relative to variable i. This means that $f(i,j)$ stands for the importance of lags of other variables relative to the importance of own lags of the dependent variable in each equation.

The literature has shown that a reasonable procedure is to set the parameter γ for overall tightness to something in the order of 0.1 or 0.2. Tightness controls the importance of a variables own lags relative to the lags of the other variables in the system. Setting it too high eliminates the Bayesian part of the system. On the other hand, setting the value for tightness too low forces the own lags of a variable too close to the prior mean of the lags of the other variables. We stick to the value of 0.2 on overall tightness.

Regarding the tightness of own lags, we set the decay factor for own lags to $g(l)$ to 1.5.

Setting the value for the tightness of other lags $f(i,j)$ to 1.0 gives a standard VAR. On the other hand, setting this value very low gives little weight to interaction of any one variable with the other variables' lags and, therefore, eliminates the 'vector' part of the VAR and yields an univariate model. This might be a feasible option if single equations in the VAR will turn out to be univariate autoregressions rather than equations interacting with the variables in the system. We set the importance of lags of other variables relative to the own lags of the dependent variable initially to 0.5. However, the bond supply equation turns out to come close to an univariate auto-regressive process in our Italian data. None of the other lags have significant t-statistics in this equation. Furthermore, in the equations for the CPI and industrial production only few lags of the real return are significant. Therefore, it seems to be reasonable to give a lower weight to the lags of the other variables. We find that the range of 0.3 to 0.4 does fine in the sense that impulse response functions are smoother and deliver more reasonable results than a higher value would do. Taking care of the fact that the bond supply, the consumer price index and industrial production come near to univariate auto-regressions we, therefore, set the value for the importance of other lags to 0.35. For countries other than Italy we choose the same values for the hyperparameters.

8.2 The Results of the BVAR

8.2.1 Italy

The identification scheme for the structural BVAR is the one elaborated on in system (14).

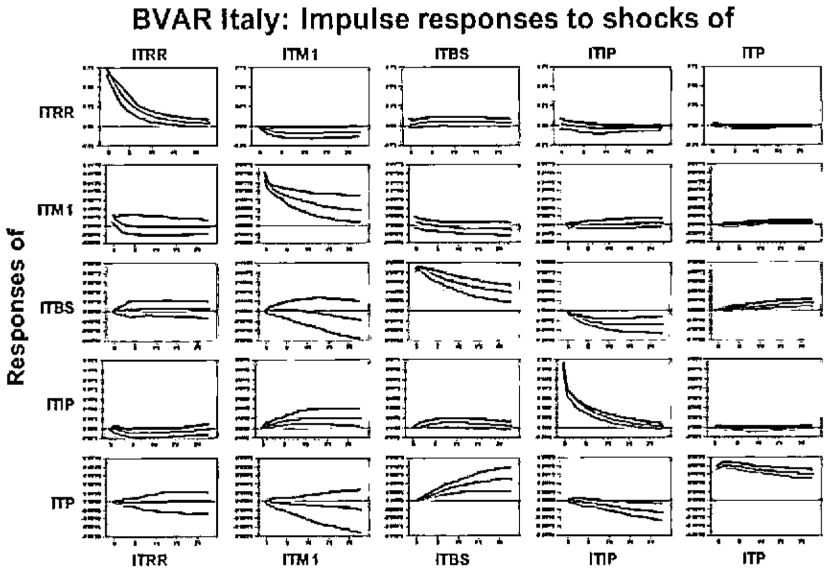
Real ex-post return of three-month treasury bills As displayed in figure 8, the Italian real return reacts positively to shocks of the bond supply as has already been indicated in the first structural model above. Regarding the magnitude of this liquidity effect, a one percentage point shock in the bond supply leads to an immediate six basis point increase in the three-month real return and to a nine basis point increase after five periods. After two years the ex-post real return still stays eight basis points up compared to its pre-shock level. These increases in the real return might be considered as being very small but certainly they are worthwhile to study. With government securities outstanding for amounts of between 255 billion euro in Belgium and 1165 billion euro in Italy differences of only very few basis points become relevant for the costs of public debt financing.

The real return shows as well the expected negative reaction to $M1$. A one percentage point monetary shock lowers the real return by twelve basis points after six periods. Although the effects of $M1$ shocks on the fluctuations in the real short-term rate seem to be slightly bigger than the effects of shocks in the bond supply, bond supply shocks put an upward pressure on real returns and contribute as well considerably to interest rate fluctuations as can be seen from the variance decomposition.

M1 Money initially shows a very subdued positive response to the interest rate shock but then declines after four periods. Two years after the shock money is still slightly lower than prior to the shock.

Bond supply The bond supply reacts positively to the real interest rate and to the consumer price index but it reacts negatively to industrial production.

Figure 8: BVAR Italy. Sample June 1985 - December 1998.



Industrial production Industrial production reacts negatively to the interest rate shock and positively to monetary innovations but the magnitude of the two effects is only marginal.

Consumer prices Consumer prices show hardly a reaction to the interest rate but a very persistent positive reaction to the bond supply. Consumer prices display a negative reaction to shocks in industrial production instead.

According to the variance decomposition in table 10 the bond supply makes

Table 10: Structural Variance Decomposition of the Italian Real Return

Period	S.E.	ITRR	ITM1	ITBS	ITIP	ITP
1	0.6253	44.8	53.2	0.0	0.3	1.7
8	1.0026	34.6	58.8	4.5	0.2	2.0
16	1.0882	30.0	59.0	8.7	0.2	2.2
24	1.1305	28.9	58.1	10.5	0.2	2.3

some contribution to the variance of the real return in Italy which increases with the forecast horizon while the immediate major contributions come from $M1$ and the real return itself.

Regarding the computation of the confidence bands for the BVAR, Sims and Zha [28] illustrate that the construction of classical confidence intervals for impulse responses in dynamic multivariate linear models is problematic because impulse responses are highly nonlinear functions of the underlying autoregressive reduced form parameters. Classical confidence bands restrict themselves to probability statements based on pre-observation probability distributions while the confidence bands for overidentified Bayesian VARs suggested by Sims and Zha [28] use probability statements about the parameters conditional on the observed data. Classical confidence regions mix information about parameter location with information about model fit and can be misleading. Error bands that illustrate the statistical reliability of estimated impulse responses should be based on the shape of the likelihood function. Therefore, we are using a RATS code from the Estima website which is based on Sims and Zha [28] in order to calculate the confidence intervals.

8.2.2 Germany

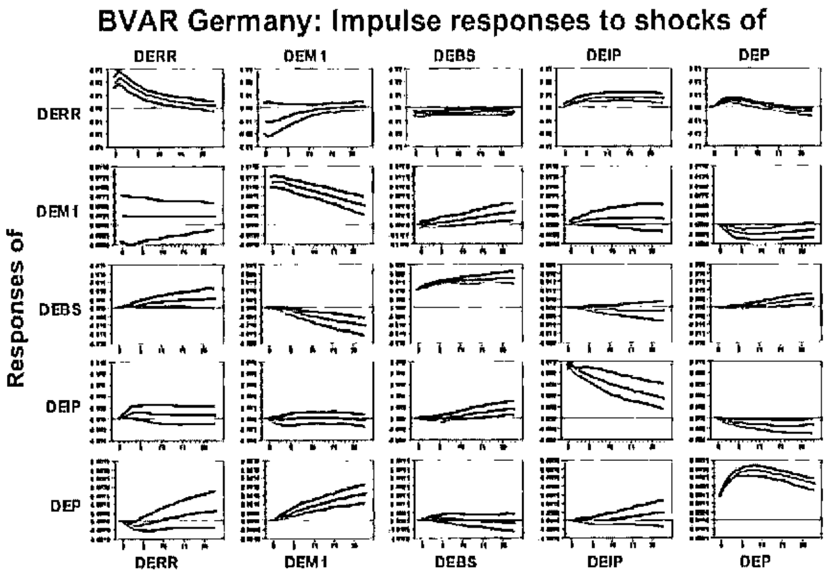
The structural model for Germany, like the ones for France, Belgium and Spain, is identical to the model (14) for Italy.

Real ex-post return of three-month treasury bills The response of the three-month ex-post real return to its own shock in figure 9 is much lower than in Italy. The three-month real return displays a negative response of about only one basis point to $M1$ shocks. In contrast to Italy the real return reacts negatively to bond supply shocks. Furthermore, the real return shows positive reactions to industrial production and, initially, to consumer prices.

M1 Money responds positively to interest rate shocks, bond supply shocks and to shocks in industrial production. $M1$ reacts instead negatively to shocks in the CPI.

Bond supply The bond supply reacts positively to the real return and to the consumer price index. It reacts negatively to $M1$ and industrial production.

Figure 9: BVAR Germany. Sample July 1980 - December 2002



Industrial production Industrial production hardly reacts to interest rate shocks and to the other shocks but shows a negative response to consumer prices.

Consumer prices Consumer prices display the correct negative reaction to the interest rate only initially and then bounce back already after six to seven periods. They show a strong positive reaction to M1 and a positive and lagged reaction to industrial production.

Table 11 shows that for Germany the bond supply contributes between two and five percent to the forecast error variance of the real return. The contribution of the bond supply and of $M1$ to the variance decomposition is much less pronounced than for the case of Italy. In contrast to Italy, the real returns evolve much more exogenously in the case of Germany even though industrial production and the CPI make an important contribution to the forecast error variance. Striking is the high contribution of consumer prices at all horizons. This reflects probably the high commitment of the Bundesbank to anti-inflationary monetary

Table 11: Structural Variance Decomposition of the German Real Return

Period	S.E.	DERR	DEM1	DEBS	DEIP	DEP
1	0.0002	69.0	0.6	2.0	0.3	28.2
8	0.0006	48.4	1.5	3.8	6.2	40.1
16	0.0007	42.9	2.8	5.1	13.6	35.6
24	0.0007	40.2	4.3	5.4	15.7	34.4

policy.

8.2.3 France

Real ex-post return of three-month treasury bills The French three-month ex-post real return in figure 10 remains practically unchanged when shocked by M1. There is no evidence for a liquidity effect. In contrast to Italy, the real return reacts negatively to shocks in the bond supply. The real return hardly reacts to the other shocks.

M1 Money shows initially a slightly negative response to interest rate shocks and reacts positively to industrial production shocks. It reacts negatively to bond supply shocks and to CPI shocks.

Bond supply The bond supply reacts negatively to the real return and to $M1$. It reacts positively to industrial production and to the CPI.

Industrial production Industrial production displays a slightly negative reaction to interest rate shocks and a positive reaction to money with strong persistence. It reacts positively to the bond supply.

Consumer prices Consumer prices show the expected negative reaction to the interest rate and display a negative reaction to $M1$, too. They show a strong positive reaction to the bond supply and a negative reaction to industrial production.

In France the bond supply constitutes a sizeable component of the forecast error variance of the real return. The variance decomposition in table 12 states that the bond supply contributes roughly between two and twenty-four percent to the variance of the real return at different horizons. In analogy to Germany

Figure 10: BVAR France. Sample February 1980 - November 1998.

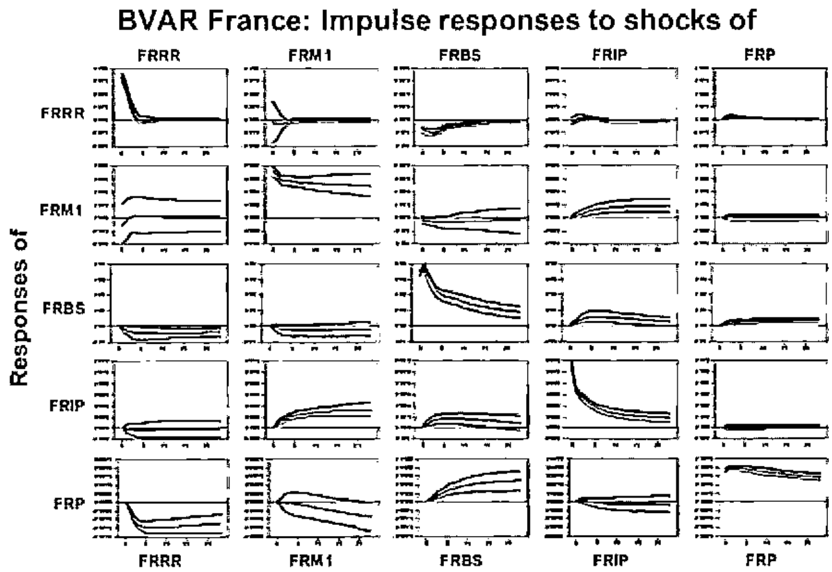


Table 12: Structural Variance Decomposition of the French Real Return

Period	S.E.	FRRR	FRM1	FRBS	FRIP	FRP
1	0.0009	58.0	5.2	2.2	0.1	34.6
8	0.0013	43.5	5.4	20.6	1.1	29.4
16	0.0013	40.5	6.8	23.4	1.3	28.0
24	0.0013	39.2	8.3	23.5	1.6	27.3

and in contrast to Italy, the French bond supply and consumer prices are an important source of fluctuations in real short-term rates while $M1$ does not contribute much to the variance in real short-term rates.

8.2.4 Belgium

Real ex-post return of three-month treasury bills Like in the case of Italy, we find evidence for the liquidity effect in the bond market for Belgium (see figure 11 for the impulse responses). The real ex-post return of Belgian three-month treasury bills increases slightly in reaction to shocks to the Belgian bond supply. Unitary shocks to the bond supply push the real return immediately nearly ten basis points higher. After two years the real return still stays three basis points higher compared to its pre-shock level. Representing the classical liquidity effect, money M1 reacts negatively to real return shocks. Concerning the other shocks, the Belgian real return hardly responds to industrial production shocks and to consumer price shocks.

M1 Money initially slightly declines due to the interest rate shock but then becomes positive already after one period. $M1$ reacts positively to the industrial production shock and to consumer prices.

Bond supply The bond supply reacts positively to the real return and to the CPI but reacts negatively to industrial production.

Industrial production Industrial production reacts positively to the real interest rate and the bond supply but negatively to $M1$.

Consumer prices Consumer prices show the correct negative reaction to the interest rate shock. After two years following a real interest rate shock prices are still considerably down. Consumer prices show positive responses to $M1$ shocks while they show negative responses to bond supply shocks.

The results of the variance decomposition for Belgium in table 13 are roughly comparable to the Italian results. The bond supply contributes between one and ten percent to the variance of the real return at different horizons. $M1$ and the real return are the other important variables and $M1$ makes the higher contribution of the two.

Figure 11: BVAR Belgium. Sample January 1980 - December 1993.

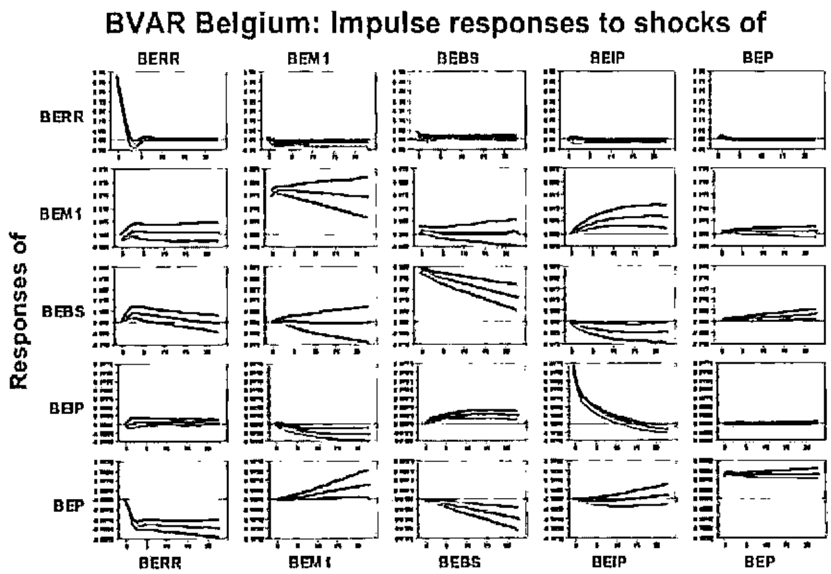
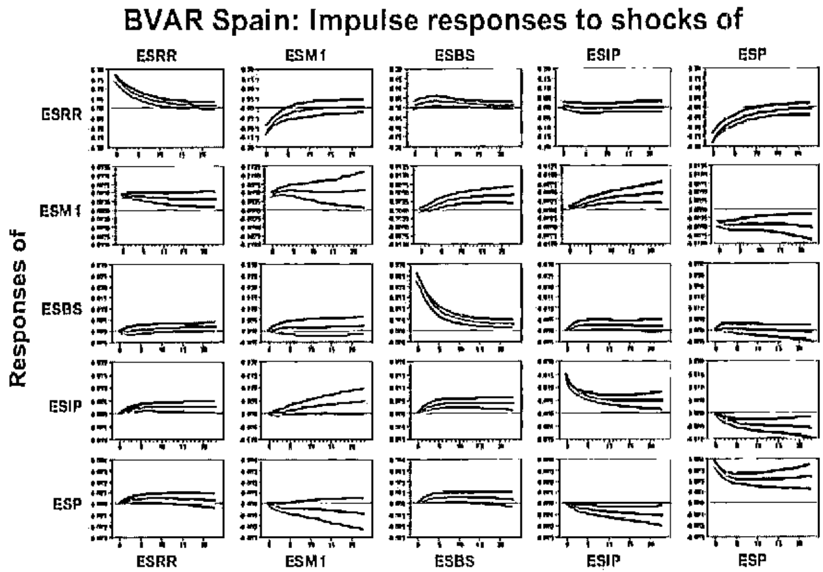


Table 13: Structural Variance Decomposition of the Belgian Real Return

Period	S.E.	BERR	BEM1	BEBS	BEIP	BEP
1	0.0046	19.3	74.0	0.7	0.0	5.9
8	0.0056	22.2	70.0	1.6	0.1	6.0
16	0.0061	23.7	65.1	4.8	0.8	5.7
24	0.0065	25.7	57.3	9.9	2.0	5.1

Figure 12: BVAR Spain. Sample March 1988 - October 1998



8.2.5 Spain

Real ex-post return of three-month treasury bills Like in the case of Italy and Belgium, we find evidence for the liquidity effect in the bond market for Spain in figure 12. The three-month ex-post real return increases by about five basis points in reaction to shocks to the Spanish bond supply. However, the traditional liquidity effect is at work, too. The real return displays a negative response to $M1$ shocks, decreasing by ten basis points in response to monetary shocks. The monetary shocks are, however, slightly less persistent than bond supply shocks. Two years after a bond supply shock the real return is still slightly above baseline while in the case of the monetary shock the real return has already returned to baseline. Besides, the real return shows a strong negative reaction to consumer price shocks indicating that CPI shocks come as a surprise to bond investors.

M1 Money displays the wrong reaction to interest rate shocks and the reaction has a high persistence. It reacts positively to the bond supply and

Table 14: Structural Variance Decomposition of the Spanish Real Return

Period	S.E.	ESRR	ESM1	ESBS	ESIP	ESP
1	0.0021	36.4	20.4	0.2	0.3	42.6
8	0.0039	38.8	13.7	7.1	1.5	38.9
16	0.0041	37.3	12.6	11.3	2.1	36.8
24	0.0042	35.8	13.3	13.2	2.4	35.3

industrial production, too, while it displays a negative reaction to consumer prices.

Bond supply The bond supply reacts slightly positively to the real return, to $M1$ and to industrial production. It reacts negatively to the price level.

Industrial production Industrial production reacts positively to the real return, $M1$ and the bond supply. It displays a negative response to price shocks.

Consumer prices Consumer prices show the wrong reaction to the interest rate. The price puzzle manifests itself in a slight increase of the consumer price index following a shock to the real return. The price level displays a negative reaction to $M1$ and industrial production shocks.

In Spain the price level plays a very important role for the forecast error variance of the real return in table 14. The same can be said for the bond supply which at longer horizons makes the same contribution to the variance of the real return like $M1$.

8.3 Comparison of BVAR and Classical VAR

For Italy the impulse responses of the BVAR and the classical VAR are similar but some shocks have very different magnitudes. The own shock of the interest rate in the BVAR in the first period is 75 bp, while it is less than ten basis points in the classical VAR. While the shock of M1 on the real return seems to be similar in both models, the shock of the CPI on the real return seems to be negative in the BVAR but positive in the classical VAR. The negative sign of the CPI shock on real returns in the BVAR seems to be more convincing if CPI shocks really come as a surprise to investors. M1 reacts initially positively to shocks in the real return but then declines after some periods in the BVAR. In the classical VAR, instead, M1 still remains slightly positive two years after an interest rate shock. The own shock of M1 seems to be too low in the classical VAR as well. For Italy the BVAR seems to deliver more convincing results in describing the dynamics of the data.

Regarding Germany, the real return's responses to the shocks seem to be better represented by the BVAR model. The reaction of M1 to the real return, instead, seems to be correct in the classical VAR while it seems to be wrong in the BVAR. The other impulse responses are quite similar or have quite large error bands so that it would be difficult to reject a null hypothesis that they are identical. Two exceptions are the shocks of M1 on M1 and on the CPI which seem to be more consistent in the BVAR.

In France the BVAR features the same problem regarding the reaction of M1 to the real return like in Germany. The classical VAR, instead, displays a negative reaction of M1 to the real return. Besides, the classical VAR suggests a positive effect of bond supply shocks to the real return while the BVAR suggests a negative relationship between the two. Another contrasting result of the two models is the reaction of industrial production to the real return which is negative in the classical model but practically zero in the BVAR.

In Belgium both the classical VAR and the BVAR suggest that shocks to the bond supply lead to higher real returns. The two models seem to be in accordance regarding the shock of M1 on the real return and the inverse shock of the real return on M1, too. The major differences between the impulse responses of the two models can be observed regarding the reactions of industrial production and the price level. In the classical VAR industrial production reacts negatively to the real return and positively to M1 and the price level reacts negatively to the real return. The BVAR seems to get all these three responses wrong.

For Spain the BVAR seems to get the reaction of the real return to M1 and to the CPI right - both are negative - while the classical VAR probably gets the wrong impulse responses. The positive reaction of the Spanish three-month real return to bond supply shocks is more marked in the BVAR than in the classical VAR. In contrast the classical VAR gets the negative reactions of M1 and of the CPI to real returns right, while the BVAR does not.

9 Conclusions

We study the effects of bond supply shocks on the real ex-post return of three month treasury bills in a structural VAR and a Bayesian VAR framework for the five euro area economies with the biggest amount of government securities outstanding. Namely, these are Belgium, France, Germany, Italy and Spain. Besides variation in monetary aggregates, variation in the bond supply should help to explain movements in nominal and in real short-term interest rates at high frequencies. Even more so, because historically the empirical literature gives no unanimous verdict on the existence, the magnitude and the sign of the traditional liquidity effect of monetary innovations. This leaves a vacuum in the explanation of short-term interest rate movements which can be partly filled up by the bond supply risk. As indicated by Jovanovic and Rousseau [15], the interest rate being the rate of return on bonds rather than the rate of return on money, one should look at the bond market rather than on monetary aggregates in order to find explanations for interest rate fluctuations.

In order to capture the dynamics between the real return on three-month treasury bills, the monetary aggregate $M1$ and the bond supply we employ a semi-structural VAR. Imposing contemporaneous restrictions on the interaction of the variables, the VAR identifies the structural shocks of the bond supply, of a monetary aggregate and of the real ex-post return on three-month treasury bills. On account of the well-known overparametrisation problems of classical VARs we estimate the structural models for our five euro area economies in a Bayesian VAR framework as well.

There seems to be a tendency that in countries in which treasury bills traditionally form a substantial part of the total amount of government securities outstanding bond supply shocks lead to higher real returns. In Belgium, Italy and Spain treasury bills represent a relatively high share of total treasury securities outstanding[4]. The same is true for the US, where the government debt agency gives a great weight to treasury bills instead of long-term bonds in the maturity structure of total government securities outstanding and where Jovanovic and Rousseau [15] find that bond supply shocks raise real ex-post three-month returns.

Our BVAR shows that in Italy a bond supply shock of one percentage point

[4]According to the ECB Securities Issues Statistics as of June 2003 the proportions of general government securities with maturities of up to one year on the total amount of general government securities were 13.4% for Belgium, 12.7% for France, 12.3% for Spain, 12.2% for Italy and 3.3% for Germany.

pushes up the three-month real return immediately by six basis points. In Belgium the real return increases by thirteen and in Spain by ten basis points in response to a bond supply shock. Two years after the bond supply shock real returns are still up by three basis points in Belgium, eight basis points in Italy and even ten basis points in Spain. In Belgium and in Spain the responses of the real returns to the bond supply shocks are more persistent than the responses to the monetary shocks even if, at least at short horizons, the magnitude of the monetary shocks is higher. At long horizons, however, the magnitude of the bond supply shocks is higher, implying that a contraction in $M1$ has a weaker upward effect on the real return than an expansion in the bond supply of the same magnitude. In Belgium the monetary shock initially even has the wrong sign. Only in Italy the monetary shock is stronger than the bond supply shock at all horizons and the monetary shock displays the higher persistence as well. As reported by Jovanovic and Rousseau [15], this is in contrast to the US where in some sub-samples the bond supply has a stronger effect on interest rates than $M1$ has.

A notable exception might be France. In France treasury bills make up a considerable share of total government securities outstanding, too. However, we find no evidence that bond supply shocks push up the three-month ex-post real return in France. In France and in Germany bond supply shocks seem to lower real returns while, strangely enough, monetary shocks can not explain short-run interest rate fluctuations because the two countries experience the liquidity puzzle. Shocks in $M1$ lead to higher interest rates. Certainly the liquidity effect in the bond market, even with the wrong sign, has still important implications for these two countries because in both countries the liquidity puzzle is accompanied by bond supply shocks whose effects - though displaying the wrong sign - have a bigger magnitude on real returns than money supply shocks have.

Although we find that government bond supply shocks raise real returns in Belgium, Italy and Spain and that, at medium horizons of about two years, bond supply shocks have bigger effects on real short-term interest rates than money supply shocks in all of our countries with the exception of Italy, we can not yet make out a pattern to group countries according to the sign of the traditional liquidity effect, the sign of the liquidity effect in the government bond market or according to the relative magnitudes and the persistence of these two effects.

These results have displayed the potential of the liquidity effect in the euro area government bond market in explaining movements in ex-post three-month real returns in the five countries analysed. One possible drawback of our work might be that, due to the scarcity in the availability of data, we use the shocks to the total amount of government bonds outstanding instead of the shocks to the net bond supply which would exclude the amounts of securities held by the respective central banks. To our understanding, excluding the amounts in the portfolios of the central banks and, thus, using only the amounts actually being available for trading to the public should, however, further strengthen the role played by the liquidity effect in the bond supply in explaining short-term interest rate movements.

Therefore, in future studies it would be desirable to prove that using the net bond supply will strengthen the relationship between the liquidity effect in the euro area government bond market and euro short-term rate fluctuations.

Furthermore, it would be interesting to analyse whether and how the upward pressure, which shocks in the bond supply exert on real returns in some countries is related to the maturity structure of the government securities of these countries and how it is related to the volatility in the government bond supply of the respective countries.

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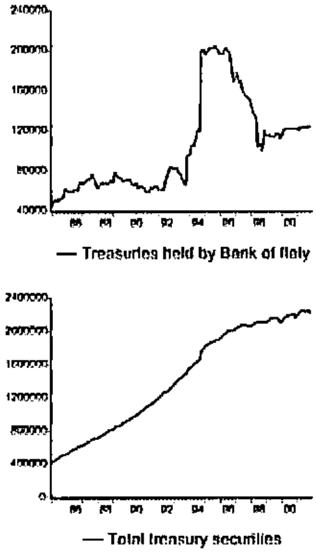
Appendix A: Italian net bond supply

In the above analysis of the effects of bond supply risk on ex-post real returns we use the total amount of government securities outstanding rather than the

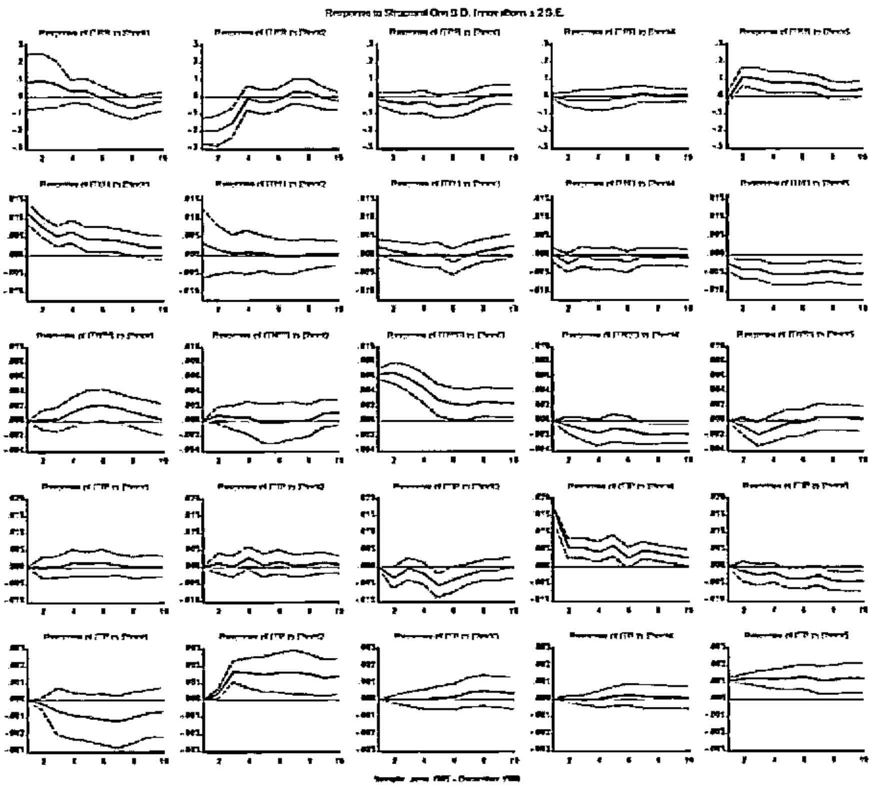
Figure 13: Total treasury securities and treasury securities held by the Bank of Italy in Italian Lira.



amount which is actually hold by private agents as an input to measure the variation of the bond supply. Therefore, we do not account for the amount of treasury securities which is held by the central banks and, thus, does not actually make part of the bond supply available for trading on private markets. This implies that if the share of securities held by the central bank varied by means of frequent and substantial open market operations the bond supply per se could differ very much from the amount of total government securities which we are using as a proxy for the bond supply in my analysis. In this case the impulse response analysis of the effects of innovations in M1 and in the bond supply on the ex-post real return in Italy represented in figure 7 could be highly misleading. Figure 13 displays for the case of Italy the total amount of treasury securities outstanding and the amount of treasury securities held in the portfolio of the Bank of Italy. The share of the securities held by the central bank is particularly volatile in the mid-1990s.

In order to control for the volatility in the central bank holdings of government securities We include the net bond supply, defined by the difference between total government securities outstanding and securities held by the cen-

Figure 14: Structural VAR including Italian net bond supply



tral bank, instead of the total amount of treasury securities into system 14. The resulting impulse responses are displayed in figure 14. Qualitatively, the impulse responses confirm the results obtained using the total amount of treasuries represented by figure 7. Even quantitatively there are only minor differences in the impulse responses.

The Risk Premium of Government Bonds and the Homogeneity of Monetary Policy Transmission in the Euro Area

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Abstract

Since the introduction of the euro, a broad strand of literature concerning the homogenisation of the monetary policy transmission process across euro area countries has unfold. This literature has largely ignored the contribution of the collapse of the country risk premia during the build-up to monetary union. We show that risk premia-adjusted yield spreads between ten-year government bonds and three-month treasury bills, which are based on the expectations hypothesis of the term structure, display roughly the same reactions to monetary policy shocks before and during EMU in euro area countries. This is in contrast to actual yield spreads which clearly change their behaviour with respect to monetary policy shocks between the two sub-periods. The monetary policy transmission mechanism has most likely become more homogeneous since the introduction of the single currency but this has probably mainly been due to the collapse in the country risk premia.

Keywords : risk premium, monetary policy transmission, euro area
JEL classification : E43, E52

1 Introduction

It has been widely argued that monetary policy transmission in the Economic and Monetary Union (EMU) countries has become more homogeneous after 1999. It is, however, somewhat problematic to establish a causal relationship between the advent of the single currency and increased homogeneity in monetary transmission. The monetary transmission process is a very broad and complex concept, which is difficult to pin down in simple economic models. Very simplified, the monetary transmission mechanism can be described in the following way. Initially, financial market interest rates react to official interest rates set by central banks. Second, market rates trigger an adjustment in asset prices, exchange rates and agents' inflationary expectations. Third, changes in asset prices and expectations cause spending behaviour in the household and corporate sectors to adjust. Finally, changes in aggregate demand bring about changes in output and in inflation.

Why could monetary policy transmission have become more homogeneous in the countries of the euro area? The single monetary policy brings along with it some fundamental transformations of the key links in the monetary transmission process such as a single money market, a common yield curve and a highly integrated debt securities market.

Monetary transmission being a very wide field, we pick up just a few but very important aspects which merit further and deeper analysis. In particular, we analyse those links of monetary transmission in which changes due to EMU are likely to occur rapidly, namely the banking sector and financial markets.

Analysing exactly these linkages, studies on monetary transmission like, e.g., Angeloni and Ehrmann [1], find evidence in favour of increased homogeneity in monetary policy transmission since the introduction of the euro in 1999.

The literature on the dynamics of the homogeneity of monetary transmission in the euro area usually neglects the effects of one very important feature of monetary unification - the collapse of the risk premia prior to the introduction of the euro in 1999. This is where the contribution of the present study kicks in.

Increased homogeneity in monetary transmission usually goes hand in hand with increased interest rate co-movement and, eventually, more symmetric reaction to shocks. In this paper we look at reactions to monetary shocks rather than reactions to fiscal shocks. If reactions to monetary shocks have become more homogeneous then why so? Here, we will have a look at the risk premia of ten-year government bonds which collapse during the second half of the 1990s.

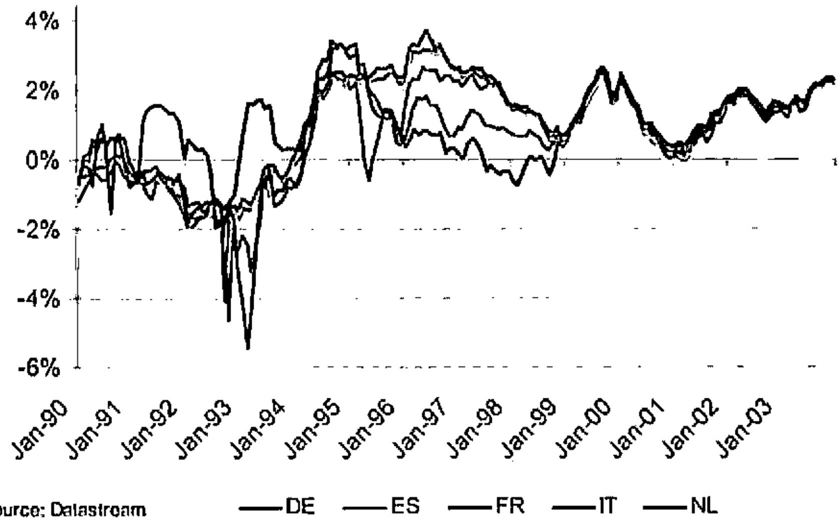
Our methodology consists of estimating a vector-autoregressive (VAR) model including the term spread of ten-year government bonds over three-month money market rates, the monthly change in the three-month money market rate and, where appropriate, the money market rate of Germany as nominal-anchor of the European Monetary System. This VAR is employed to forecast changes in future short-term interest rates in-sample. We use the forecast three-month rate changes to calculate the expectations-related component of the yield spread and compare it to the actual spread in order to extract the risk premium component. This allows us to compare the impulse responses of the actual term spread and the impulse responses of the risk-adjusted term spread, net of the risk premium component, to shocks in the three-month money market rate.

The analysis of the impulse responses of the actual term spread and of the theoretical term spread, derived from the Expectations Hypothesis of the Term Structure, to shocks in monetary policy rates further elicits the question of homogeneity of the monetary transmission process. The term spread and, more generally, the term structure represent important channels for monetary policy transmission. The monetary authorities have direct influence over short-term interest rates, while long-term rates evolve on the basis of investors' expectations about the future path of short-term interest rates and inflation. Investment, savings and output adjust in response to these expectations. Hence, the term structure transmits monetary policy to the real sector of the economy.

Since 1999 eurozone countries have been displaying term structure-derived term premia which are nearly identical and which are very much reduced in magnitude compared to the risk premia which prevailed during the financial turmoil of the exchange rate crises at the beginning of the 1990s. Controlling for these risk premia in term spread data from 1990 to 1998 we will show in section 5 that the reactions to monetary policy shocks are quite symmetric between eurozone countries even in this period of high exchange rate volatility and they are not much different to the reactions of the risk-adjusted term spread during the EMU period from 1999 to 2004. Therefore, a crucial factor in the increased homogeneity of monetary transmission seems in reality to be the collapse of the risk premia during the build up to monetary union and with the actual foundation of the monetary union in 1999. This risk premium component seems to explain the differences in the reactions to monetary policy shocks between the pre-EMU period and the EMU period up to a high degree.

Before we will derive the risk premia from the expectations theory of the term

Figure 1: Euro area countries' $(10y - 3m)$ term spreads at monthly frequency

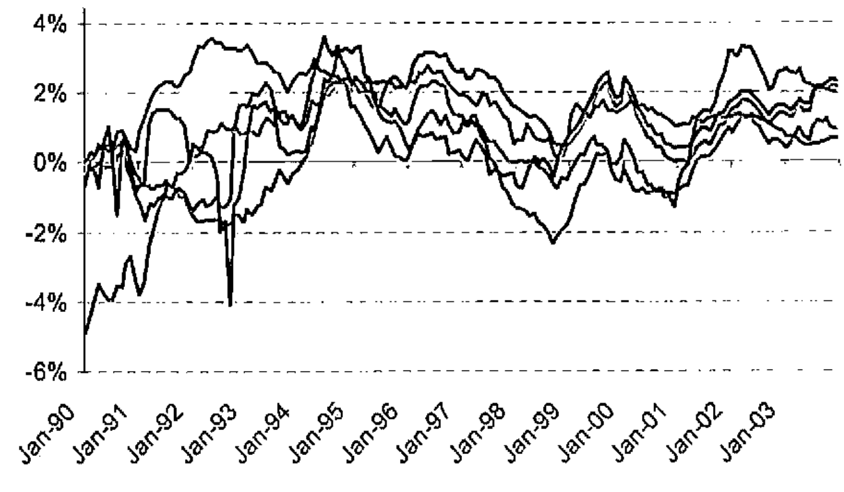


structure in later chapters, note the exemplary convergence of the term spreads between ten-year government bond yields and three-month money market rates in Germany, France, Italy, the Netherlands and Spain in figure 1. The term spreads converge in the time up to the introduction of the single currency in 1999 and only a difference of a few basis points remains between the spreads of the five countries. The same will be true for the risk premia of the five countries as we will prove below in section 4.3.2.

Clearly, the collapse of the term spread is a phenomenon which is specific to the euro area rather than being a feature of the integration of international financial markets in general, as can be seen by contrasting the German and Italian term spreads with the term spreads of some countries of a control group displayed by figure 2.

Obviously, the degree of convergence of the risk premia has important implications for some of the key links in the monetary transmission process. The collapse of the risk premia should even out the way in which economies in the euro area react to monetary policy shocks. Eurozone countries have a common term structure and a common yield curve by definition. This very trivial fact

Figure 2: Control group countries' ($10y-3m$) term spreads at monthly frequency



Source: Datastream

has been ignored or at least has not been formally analysed by the major part of studies on the homogeneity of monetary transmission in the eurozone.

Facing an identical yield curve, agents in different countries are more likely to take the same consumption and investment decisions and to form roughly the same expectations with regards to inflation. Therefore, output and prices will react more homogeneously across countries which form a monetary union. The fact that the dynamics of the term spreads of different countries converge and eventually become quite uniform makes it more likely that other links in the monetary policy transmission mechanism become more symmetric, too. Faced with practically identical term spreads and yield curves consumers in different countries have a higher propensity to take similar intertemporal decisions on consumption and savings and to build similar inflationary expectations and this translates into increased homogeneity of the dynamics of aggregate demand, prices and output in the different countries.

This paper is structured in 6 sections. In section 2 we give a short overview on the literature on monetary transmission to put the scope of our study into context. Section 3 analyses the interest rate channel of monetary policy transmission. In section 4 we derive the term premium from the Expectations Hypothesis of the Term Structure and in section 5 we analyse the implications the term premium has for the degree of homogeneity of the monetary policy transmission process in some of the eurozone countries. Finally, in section 6 we draw the conclusions from our established evidence.

In appendix A, we analyse as well the dynamics of the bank-lending channel of monetary policy transmission which yields some evidence for increased homogeneity of the monetary transmission process across EMU countries.

2 Literature on Monetary Policy Transmission in the Euro Area

What are the effects of monetary policy changes on prices and output in the euro area? How are these effects channelled? How do effects of monetary policy differ over time, across countries or across individual sectors of the economy?

A wide range of empirical literature on monetary policy transmission in the euro area has addressed these questions. Generally speaking, monetary policy has temporary effects on euro area output and, initially muted, but highly persistent effects on consumer prices.

Angeloni and Ehrmann [1], find evidence in favour of increased homogeneity in monetary policy transmission since the introduction of the euro in 1999. They analyse the banking channel, the asset market channel and the interest rate channel of the monetary transmission mechanism. They neglect, however, the fact that the risk premia of eurozone countries collapse, and, hence, they do not analyse homogeneity before and after monetary union in a risk-adjusted interest rate setting.

Van Els et al. [35] carry out a much wider study of transmission which takes care as well of the eventual effects of monetary policy on prices and output and find considerable homogeneity in the responses of prices and output to monetary policy shocks across euro area countries.

Instead, the ECB in its monthly bulletin in October 2002 reports that 'The empirical evidence does not suggest that there are systematic differences between countries in policy transmission that are robust across different studies and methodologies'. This does, however, not mean that monetary transmission works in the same way in all the countries. A main drawback of the employed models is that generally they are not robust across different methodologies and, therefore, they are not capable of providing clear statistical evidence of differences in national transmission mechanisms as Berben et al. [5] point out.

Berben et al. [5] analyse differences in transmission of euro area monetary policy in large-scale central bank models in order to find out, to what degree these differences are due to differences in the underlying economics and to what degree they are due to methodological differences and modelling strategies. They find that in most cases, significant cross-country differences seem to be reasonable and meaningful when put into context with the characteristics of the underlying economies. Financial structures and the fiscal policy framework may explain in part the heterogeneity in responses of countries to monetary policy shocks. Other structural differences in the economies which are shown to be important are differences in labour markets and consumption and investment behaviour. However, differing model strategies or model features can play a role sometimes and cast doubt on the validity of models.

The studies of the ECB's Working Group on Econometric Modelling (WGEM) underline the different results achieved by differing methodologies when comparing reactions of output and the price level from the Area Wide Model (AWM) and from the national central banks' country models (NCB). Both models suggest that output reacts temporarily to a monetary policy shock but the magnitude of the response in the AWM is double the magnitude in the NCB model in the second and the third year after the shock.

Confronting instead the WGEM results with the VAR evidence yields only small cross-country divergence in price responses but substantial differences in output responses.

Concerning studies on the degree of homogeneity of monetary policy transmission, the main caveat is that monetary transmission depends up to a large extent on the monetary policy regime itself. Therefore, the change in the policy regime due to the introduction of the euro in 1999 makes it very difficult to draw any inference from studies which report a possible increase of homogeneity in the monetary transmission process. Results of studies on the magnitude of policy effects often are not robust across different methodologies. Differences in the estimated impact of monetary policy on output and prices across countries do not tend to be robust across methodologies, data and models.

3 The Interest Rate Channel of Monetary Policy Transmission

The financial markets channel is somewhat broader and, thus, its interdependencies are more complex than for example the ones of the bank lending channel. Transmission through financial markets can take place in numerous different ways. Therefore, we concentrate exclusively on analysing the interest rate channel effect. The interest rate channel effect is the traditional Keynesian effect whereby monetary policy is transmitted, through liquidity and expectations effects, to the structure of nominal and real interest rates, and then, indirectly, to investment and consumption plans of non-financial firms and households.[1]

By arbitrage, nominal long-term rates in the euro are converged after the elimination of exchange rate risk in 1999, apart from small remaining yield spreads in the order of 10 to 20 basis points versus the German benchmark bonds. These spreads prevail on the grounds of liquidity and risk premia.

If the interest rate channel has become more homogeneous in the Euro area after 1999, then it must be true that rates in real terms, too, must converge to a certain degree across all countries or must at least show a high degree of co-movement between all the member states. It remains to be seen, whether greater homogeneity can be ascribed to signal-quality or market structure effects as Angeloni and Ehrmann [1] point out. Our hypothesis is that the collapse of the risk premia is the main factor influencing higher co-movement of interest rates. Near-convergence in real rates could have been caused largely by the by and large elimination of the risk premia due to the removal of exchange rates in 1999. However, we analyse the implications of a more integrated banking market and of the banking channel of monetary transmission in appendix A, too.

In order to shed some light on the question whether homogeneity in monetary transmission has increased, we construct risk premium-adjusted real interest rates and term spreads and analyse their dynamic behaviour between 1990 and 2004. The expectations hypothesis of the term structure provides us with a measure of the risk premium which then is applied to adjust interest rates on ten-year government bonds and term spreads for risk.

Angeloni and Ehrmann [1] conclude that converging real rates could be a proof of a higher degree of homogeneity in the monetary transmission process.

[1] Angeloni and Ehrmann [1]

Table 1: Correlations of monthly three-month real interest rate changes before and after monetary unification

	DE	ES	FR	IT	JP	NL	UK	US
DE	1	0.20	0.60	0.55	0.18	0.28	−0.05	0.37
ES	0.17	1	0.46	0.12	0.15	−0.02	0.23	0.40
FR	0.25	0.23	1	0.47	0.03	0.14	0.18	0.48
IT	0.04	0.45	0.06	1	0.35	0.11	0.04	0.50
JP	0.09	−0.03	−0.10	0.01	1	0.16	0.10	0.21
NL	0.29	0.22	0.38	0.17	0.01	1	0.06	0.05
UK	0.03	−0.06	0.09	0.09	0.09	0.20	1	0.32
US	0.07	−0.12	−0.11	−0.10	0.07	0.04	−0.18	1

upper triangle: correl. of monthly real short-rate changes Jan'99-Jan'04
lower triangle: correl. of monthly real short-rate changes Jan'90-Dec'98
correlations between eurozone countries are reported in bold

In order to compare pre-EMU and post-EMU evidence they chose to measure the co-movement of real interest rates at different maturities across countries. If real interest rates in EMU countries co-move to a high degree this could imply that the effects of monetary policy on interest rates are very similar in these countries.

We retrieve data for three-month money market rates and CPI inflation from DataStream in order to calculate short-term real interest rates for five big euro area economies and three countries of a control group. For our data the correlations of monthly real interest rate changes between the pre-EMU period (1990-1998) and the EMU period (1999-2004) increase between all of the EMU countries with the notable exception of the Netherlands as can be seen from table 1. The rates being used to calculate the correlations are three-month money market rates adjusted for the annual consumer price inflation in each country.

The correlation of real short-rate changes did not only increase among eurozone countries themselves but it increased as well between eurozone countries and the three economies of our control group. In particular, as displayed by the last column in table 1, the co-movement between the eurozone countries' short-rates and US short rates increased much more than the co-movement between the eurozone countries' short-rates among themselves. Therefore, if higher interest rate co-movement is considered a proxy for higher homogeneity in mon-

Table 2: Correlations of monthly ten-year real interest rate changes before and after monetary unification

	DE	ES	FR	IT	JP	NL	UK	US
DE	**1**	**0.99**	**1.00**	**0.99**	0.59	**1.00**	0.90	0.76
ES	**0.93**	**1**	**1.00**	**1.00**	0.64	**1.00**	0.89	0.80
FR	**0.98**	**0.96**	**1**	**0.99**	0.61	**1.00**	0.90	0.78
IT	**0.87**	**0.92**	**0.88**	**1**	0.60	**0.99**	0.87	0.76
JP	0.96	0.90	0.95	0.77	1	0.61	0.56	0.89
NL	**1.00**	**0.92**	**0.99**	**0.86**	0.95	**1**	0.90	0.78
UK	0.96	0.93	0.96	0.81	0.94	0.96	1	0.75
US	0.90	0.81	0.89	0.66	0.89	0.90	0.94	1

upper triangle: correl. of monthly 10y real rate chg. Jan'99-Jan'04
lower triangle: correl. of monthly 10y real rate chg. Jan'90-Dec'98
correlations between eurozone countries are reported in bold

etary transmission, the question, whether the monetary transmission process has really become more homogeneous because of EMU or whether the higher co-movement of short-term real rates among eurozone countries is in reality a part of the bigger picture of more integrated global financial markets, has to be raised. This is a question which is beyond the scope of this paper because instead of analysing the increasing financial integration of the world economy our motivation is to study possible structural changes in the reaction to monetary policy shocks between 1990 and 2004.

Is the higher co-movement in the eurozone due to the signal quality effect, the market structure effect or due to the collapse of the risk premium? We will address and investigate this issue in an analysis of the co-movement of risk free interest rates and in an analysis of the reaction of interest rates to monetary policy shocks in section 5.

Regarding the longer maturities, we look at the real yields of ten-year government bonds and their evolution over time. Again we take the data for these ten-year constant maturity bonds from DataStream. The correlations of ten-year real rate changes have been increasing mainly between eurozone countries themselves as displayed in table 2. The co-movement of long-term interest rates between EMU countries presumably increases already much earlier than 1999, during the build-up to EMU, so that the results displayed in table 2 were much stronger if one would be prepared to split the sample, let's say, already in 1995

Table 3: Correlation of monthly term spread changes before and after monetary unification

	DE	ES	FR	IT	JP	NL	UK	US
DE	1	**0.84**	**0.88**	**0.99**	0.19	**0.88**	0.76	0.67
ES	**0.51**	1	**0.97**	**0.83**	0.14	**0.72**	0.69	0.55
FR	**0.21**	**0.20**	1	**0.87**	0.24	**0.75**	0.71	0.59
IT	**0.27**	**0.20**	**−0.12**	1	0.21	**0.89**	0.77	0.66
JP	0.39	0.24	0.05	0.09	1	0.16	0.24	0.22
NL	**0.73**	**0.51**	**0.23**	**0.31**	0.30	1	0.69	0.58
UK	0.45	0.35	0.04	0.45	0.22	0.47	1	0.62
US	0.28	0.21	0.06	0.13	0.29	0.23	0.30	1

upper triangle: correl. of monthly term spread chg. Jan'99-Jan'04
lower triangle: correl. of monthly term spread chg. Jan'90-Dec'98
correlations between eurozone countries are reported in bold

rather than in 1999.

In contrast, the correlations between monthly changes in eurozone countries' real long-term interest rates with the ones of the US, the UK and Japan display a marked decrease after 1999, although, correlations of changes in nominal interest rates increase universally between the eight countries of our study after 1999.

An eliciting way of describing the dynamics of interest rate movements in EMU is to look at the term spreads between ten-year and three-month interest rates in the different countries. The term spreads contain information on future short-rate and long-rate dynamics, as well as information on the development of future output and inflation. The term spread changes in table 3 are constructed as the monthly changes of the difference between ten-year and three-month real interest rates in the respective countries.

The term spreads display dynamics very similar to the dynamics of the interest rate co-movement. As can be seen from table 3, term spread changes do not only become more homogeneous among eurozone countries but become as well more homogeneous between the eurozone countries on the one hand and the UK and the US on the other hand.

This kind of evidence has been used to suggest that the co-movement of interest rates and the monetary transmission process have become more homogeneous due to a signal quality and to a market structure effect after monetary unification, see for example Angeloni and Ehrmann [1]. From our perspective, it remains to be seen, however, in as much the common squeeze of the risk premium of eurozone countries has contributed to the increased co-movement of the eurozone countries' interest rates and term spreads and to the observed increased homogeneity of the monetary policy transmission mechanism. To this end, we will measure the risk premia inherent to ten-year interest rates and analyse the co-movement of risk-adjusted interest rates and term spreads. If risk-adjusted term spreads displayed at least as high a degree of homogeneity in interest rate dynamics than the 'raw' term spreads in table 3 do, we could infer that the higher degree of homogeneity in interest rate dynamics and in monetary transmission found for the eurozone countries by other studies is mainly due to the crunch of the risk premium rather than to signal quality or market structure effects.

4 The Risk Premium in the Euro Area Countries

4.1 The Expectations Hypothesis of the Term Structure

The Expectations Hypothesis of the Term Structure is the principal economic theory that links short term interest rates to long term interest rates. Popularised by Fisher [19] and Keynes [23], it has been widely studied by scholars, and for many economists it continues to be the way in which they think about the determination of long-term interest rates.

The term structure contains useful information on the intertemporal optimisation of economic agents. Does the slope of the yield curve predict future changes in interest rates? The answer to this question and the solution to the analysis of the co-movement of interest rates should be delivered by the expectations theory of the term structure. According to the expectations theory the expectations about future interest rates are the dominant force determining movements in long-term interest rates. If the expectations theory fails, predictable changes in excess returns of long-term interest rates must be the main influence of changes in the term structure.

The expectations hypothesis has been extensively examined in empirical work and the joint hypothesis of rational expectations and the pure expectations theory has been rejected in most studies. At least some studies of the expectations theory, including Fama [16], Campbell and Shiller [6], Mishkin [25], Tzavalis and Wickens [33] and Sutton [32] find empirical evidence in favour of some role that the expectations theory can play in forecasting future short-term interest rates. The results of Campbell and Shiller [6] and of Hardouvelis [22] are mixed, however. Campbell and Shiller [6] find an important element of truth in the expectations theory although rejecting it statistically. Hardouvelis [22] rejects the expectations theory in a sample from 1954 to 1992 for US data, identifying large deviations of long-rates from the levels predicted by theory.

Strong support for the applicability of the expectations theory is brought forward by Gerlach and Smets [20]. They study one-, three-, six- and twelve-month euro-rates for 17 countries and find positive evidence for a weak form of the expectations hypothesis in conjunction with the presence of a time-varying term premium for a broad range of countries. Sutton [32], too, intervenes for the defence of the expectations hypothesis and finds support for the expectations hypothesis for US data when Bayesian inference is used. Favero and Mosca [18] find that particularly in times of low uncertainty about monetary policy the expectations hypothesis delivers a good description of the relationship between the yields of six-month and three-month treasury bills in the US.

However, Shiller et al. [31] and Mankiw [24] point out that the term spread predicts the wrong sign for subsequent changes to the long rate. This result is particularly pronounced for the United States, the country with the most sophisticated and liquid financial markets. For the US the term structure is generally rejected for the time of the Volcker experiment between 1979 and 1982. In the time before 1979 and after 1982 there is more support for the expectations hypothesis.

Gerlach and Smets [20] suggest that the expectations theory works best in settings of fixed exchange rates such as the ERM. With constrained monetary policy short-term rates become more predictable favouring the functioning of the expectations theory. We find some support for this in the time after the ERM crisis from 1994 onwards.

The great variety in the findings of the literature partly reflect the differences in specifications, empirical approaches and data. At least for the US, the expectations hypothesis receives considerable support from studies carried out on pre-1979 data and on data between 1984 and 1990. On the whole, there seems to be a general agreement that the very short end of the yield curve displays some ability to predict changes in short-term rates, though this predictive power fades as the horizon lengthens.

Our intention not being to test the validity of the expectations hypothesis per se, but measuring the risk premium in order to take a stance on the risk premia and their likely effect on monetary transmission before and after monetary unification, we choose Campbell and Shiller's [6] method for asset price valuation and compare the theoretical price of the asset to its market price in order to derive the term premium inherent to the asset.

4.2 The Expectations Hypothesis and Present Value Models

Instead of simply rejecting the expectations hypothesis on the grounds of the statistical significance of coefficients, Campbell and Shiller [6] develop an informal evaluation of the validity of the expectations hypothesis on the basis of model fit. Numerous studies reject the expectations theory statistically. It is, however, more meaningful to evaluate the theory informally and to quantify which degree of variance in long-term rates and in short-term rates the theory is actually capable of explaining. To this end, theoretical asset prices from VAR based forecasts of fundamentals are calculated and then compared with actual asset prices. If high correlation can be found between actual and theoretical prices, the theory could be regarded as a relative success even if statistical tests rejected the model.

More formal specification tests are the variance of the error term models of the expectations theory, the ratio of the standard deviations of the theoretical and the actual spread and the slope coefficient of a regression of the theoretical spread on the actual spread. Campbell and Shiller [6] do not reject the expectations hypothesis as the actual and the theoretical spread co-move quite closely in their sample. Hence, they conclude that there is an important element of truth in the expectations theory of the term structure.

The expectations theory establishes a well-defined relationship between long-term rates and short-term rates. The current term spread between a long-term interest rate and a short-term interest rate is assumed to be the weighted average of the expected future changes in the short-term interest rate.

$$
\begin{aligned}
R_t^n &= \frac{1}{k}\sum_{i=0}^{k-1} E_t r_{t+mi}^m + tp. && (1)\\
\text{where } k &= \frac{n}{m} && (2)
\end{aligned}
$$

For a pure discount bond the long rate R_t^n should correspond to a constant tp plus the average of the current and expected future short-term rates r_{t+mi}^m up to time $(n-m)$. tp is the term premium which varies over the maturities n and m but which according to theory should not vary over time. The longer the maturity of the security, the greater is the risk of fluctuations in the value of the principal to the investor. If there were no uncertainty about the path of future interest rates, today's long-rate would equal a weighted sum of expected future short-rates without a term premium.

Regarding the formation of expectations, the expectations hypothesis makes the simple but very unrealistic assumption of perfect knowledge of future short-term rates and, hence, perfect knowledge of future long-term rates as well.

However, as future short rates are not known with certainty, risk-averse investors will require a term premium to bear this interest rate risk. Risk-averse investors prefer to lend short, unless they are offered a premium in compensation for lending long and bearing the interest rate risk of doing so. The term premium they require is a predictable excess return on the n-period long bond over the lifetime m of the short bond. Therefore, in contradiction to the pure expectations hypothesis a term premium tp has to be modelled if the goal is to analyse a weaker, but more realistic form of the expectations hypothesis.

The spread between the long-term interest rate R_t^n and the short-term rate r_t^m is a weighted average of the expected future short-term interest rate changes Δr_{t+mi} , where the weights decline over the term to maturity of the long-term bond.

$$R_t^n - r_t^m = E_t\left[\sum_{i=1}^{k-1} \frac{k-i}{k} \Delta r_{t+mi}\right] + tp \tag{3}$$

Shiller [30] points out the necessity of applying weights to future interest rate changes. Expected short-term interest rate changes in the near future should carry more weight in determining the long yield than expected short-term rate changes in the more distant future.

The term spread should predict the change in the yield of the long-term bond R_t^n during the m-period of the life of the short-term bond. The long term-bond has n periods to mature at time t. So it has $(n-m)$ periods to maturity at time $(t+m)$ at which the short bond matures. Therefore, according to the expectations theory the adjusted spread $\left(\frac{m}{n-m}\right) S_t^{(n,m)}$ between the long-term rate R_t^n and the short-term rate r_t^m can be written as:

$$\left(\frac{m}{n-m}\right) S_t^{(n,m)} = E_t R_{t+m}^{(n-m)} - R_t^n \tag{4}$$

The spread is weighted by the maturity difference between the short-term and the long-term rate $\left(\frac{m}{n-m}\right)$. This yields the predicted value of the term spread. Regressing $E_t R_{t+m}^{(n-m)} - R_t^n$ onto a constant and its predicted value $\left(\frac{m}{n-m}\right) S_t^{(n,m)}$ should yield a slope coefficient of one. According to the expectations theory, a positive term spread predicts a future increase of the long-term yield paired with even stronger increases in the short-term yields, so as to close the gap between the two. Empirically, the term spread has not been a very good indicator of the yield change of the long-term bond yield during the life-time of the short-term bond. Campbell and Shiller [6] find that when the yield spread is high, longer-term rates behave contrary to the expectations theory, whereas short-term rates behave in accordance with the theory. The empirical rejection of the expectations theory could be due to a time-varying term premium embedded in long-term interest rates. Risk-averse investors might require a term premium when lending long, instead of lending short, and the magnitude of the premium might vary over time according to differing expectations about future output, inflation and interest rates. In fact, term premia often decline at the beginning of booms and long-bond yields might decline instead of increasing in line with the expectations hypothesis.

Moving on to measure the term premium inherent to long-term interest rates we subtract the short rate r_t^m from both sides of equation 1. This yields

$$R_t^n - r_t^m = \frac{1}{k}\sum_{i=0}^{k-1} E_t r_{t+mi}^m - r_t^m + c \tag{5}$$

$$S_t^{(n,m)} = \frac{1}{k}\sum_{i=1}^{k-1}\left(\sum_{j=1}^{i} \Delta^m r_{t+jm}^m\right), \tag{6}$$

where $S_t^{(n,m)}$ is the term spread between the long-term yield R_t^n and the short-term yield r_t^m. If the expectations theory gave an accurate picture of the 'true' term strucuture of interest rates, the actual spread $S_t^{(n,m)}$ would be equal to the expected theory-based spread $E_t S_t^{(n,m)^*}$:

$$S_t^{(n,m)} = E_t S_t^{(n,m)^*}, \tag{7}$$

$$\text{where } S_t^{(n,m)^*} = \sum_{i=1}^{k-1} \left(\frac{k-i}{k}\right) \Delta^m r_{t+im}^m \qquad (8)$$

$S_t^{(n,m)^*}$ is the theoretical spread, where Δ^m stands for a change in the short-rate measured over m periods. The theoretical spread would prevail if agents had perfect foresight about future interest rates. This model can be evaluated by regressing the theoretical spread from equation 8 on a constant and on the actual spread from equation 4. In order to confirm the validity of the expectations hypothesis, the slope coefficient of this regression should be one.

An important assumption of the estimation method underlying the work of Campbell and Shiller is that forecasts of short rate changes generated from a bivariate VAR in the short rate changes and the term spread adequately represent market participants' expectations. Even more troubling might be that a constant-coefficient VAR estimated over a long sample and probably over several monetary policy regimes might not adequately capture the evolution of market participants' expectations.

4.3 VAR Approach of Present Value Models

Campbell and Shiller [6] develop what is called the VAR approach for evaluating present value models. A simple bivariate VAR is used to mimic the market participants' expectations regarding the short rate, the long rate and the slope of the yield curve. Therefore, an important assumption of this methodology is that the forecasts from a bivariate VAR adequately model the market's expectations. The VAR includes the one-period change in the m-period interest rate Δr_t, and the actual yield spread S_t. The VAR forecast proxies market expectations and the forecast short-rate changes procured from the VAR can be used in order to calculate the theoretical term spread S^* according to the expectations hypothesis. Long-run interest rates are inferred rather than estimated from short-run interest rate behaviour. The weighted average of forecast interest rate changes, i.e. the theoretical spread, can be calculated and compared with the actual spread. If the expectations theory is correct, the two will be identical. If not, their difference can be interpreted as the term premium.

Present value models for bonds, stocks and other economic variables are not free of controversy. There are many different ways to test the models but it is not clear how these tests relate. statistical rejections of the model might not contain much information on the real economic significance of the models and the variables in the models usually require transformations, like the removal of deterministic linear trends, which can bias test results.

The present value model of Campbell and Shiller [6] has been employed to study excess returns in both, bond and stock markets. Y_t is a linear function of the present discounted value of expected future y_t.

$$Y_t = \theta(1-\delta)\sum_{i=0}^{\infty}\delta^i E_t y_{t+i} + c \tag{9}$$

Applied to the expectations hypothesis of the term structure, Y_t is a long-term yield, y_{t+i} is a short-term yield, θ is a proportionality coefficient which is only of interest for stocks and therefore will be set equal to one, and c is a constant.

Subtracting θy_t from both sides of 9 and rearranging, yields

$$S_t = E_t\left[\theta\sum_{i=0}^{\infty}\delta^i \Delta y_{t+i}\right] + c \tag{10}$$

$$S_t = E_t S_t^* + c, \tag{11}$$

where

$$S_t^* = \left[\theta\sum_{i=0}^{\infty}\delta^i \Delta y_{t+i}\right] \tag{12}$$

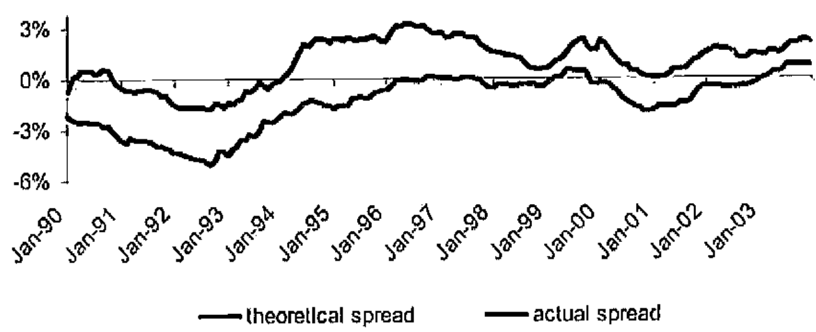
is interpreted as the theoretical spread which would prevail if the expectations hypothesis were true.

According to equation 8 from Campell and Shiller [8], we add decaying weights for short rate changes that lie farther in the future.

4.3.1 Testing the Expectations Hypothesis

The fit of the actual and the theoretical term spread We use three-month money market yield data and ten-year government bond yield data from DataStream in order to estimate a VAR for each of the eight countries of our dataset. The sample runs from 1990 to 2004 and the variables in the VAR are the term spread between ten-year bonds and three-month money market rates and, where appropriate, the German three-month money market rate to control for Germany's role as nominal anchor in the European Monetary System. The

Figure 3: Theoretical and actual term spread Germany



VAR is estimated on the full sample and the forecast short rate changes are used to calculate the risk premium according to equation 8 from above:

$$\text{where } S_t^{(n,m)^*} = \sum_{i=1}^{k-1} \left(\frac{k-i}{k}\right) \Delta^m r_{t+im}^m$$

The theoretical spread from 8 together with the actual term spread are displayed in exemplary fashion for Germany and Italy in figures 3 and 4. Compared to Sutton's [32] study on the US term structure between 1959 and 1978, there is a bigger differential between the term spreads and the theoretical spreads in Italy and Germany, indicating a higher term premium than in the US. Although there seems to be a substantial risk premium inherent to the Italian and German term spreads, the movements in term spreads can be explained to a large extent by shifts in VAR-based forecasts of future interest rate changes, i.e. the theoretical spreads. In fact the correlation coefficient between the actual and the theoretical spread is 0.87 for Germany and 0.45 for Italy.

A standard test of the expectations hypothesis is the regression of the theoretical spread on the actual spread and a constant.

As displayed by table 4, the regression of the theoretical spread on the actual spread and on a constant shows that the coefficients on the term spread are generally quite close to one and in accord with the expectations theory. On a five per cent level the Wald test fails to reject the expectations hypothesis for any

Figure 4: Theoretical and actual term spread Italy

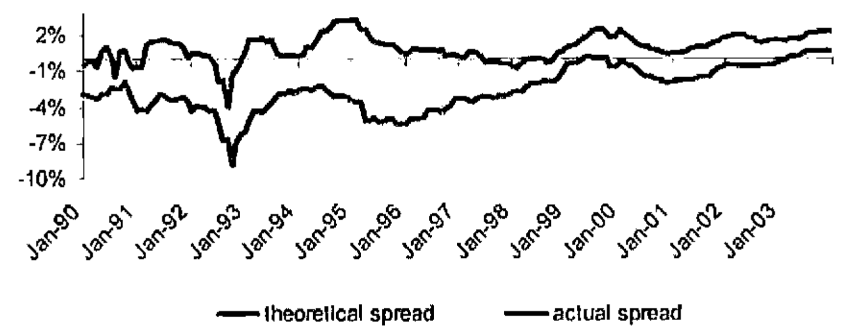


Table 4: Test of expectations hypothesis: regression of theoretical spread on term spread

	DE	ES	FR	IT	JP	NL	UK	US
const	−2.42	−3.73	−2.55	−3.53	−3.17	−2.65	−1.86	−3.61
t-stat	−18.9	−15.6	−19.5	−10.2	−9.0	−18.4	−11.2	−15.7
spread	1.00	1.18	1.03	0.86	1.38	1.04	0.94	1.15
t-stat	11.8	5.7	14.7	3.0	6.1	11.6	9.5	7.1
Wald	1.00	0.39	0.65	0.63	0.09	0.62	0.55	0.34

sample: 1990 - 2003

t-statistics are based on Newey-West autocorrelation consistent standard errors

Wald test: p value for null-hypothesis that coefficient on term spread equals one of the countries of our sample between 1990 and 2004. The negative constant is common to all countries and indicates that the theoretical spread is consistently lower than the term spread. According to the expectations hypothesis, the difference is a constant risk premium which varies between 1.86% for the UK and 3.73% for Spain.

A caveat is that specification problems call for caution in the interpretation of the results of expectations hypothesis tests because the equation linking the theoretical spread to the term spread is prone to problems with unit roots and autocorrelation in the residuals. Even if this creates problems in the interpretation of the estimation results, the high correlations of the theoretical and actual spreads (see figures 3, 4 and table 7) for many countries, nevertheless, show that there is an important element of truth in the expectations theory. As displayed in table 7 below, the correlation between the theoretical and the actual spread is around 90% for France, Germany, the Netherlands, the UK and the US. An expectations model with a time-varying risk premium with a comparatively low variance could provide a good description of the dynamics of the term structure for these countries. On the other hand, the correlation of the theoretical spread with the term spread is very low for the case of Spain and Italy indicating the statistical rejection of the expectations hypothesis for these countries. The inclusion of the time of the ERM crisis in the sample reinforces the rejection of the expectations hypothesis in these countries from both, a statistical, model-based, point of view and a more informal, model-fit, point of view.

Predicting interest rate changes We fit the following standard equation to test the expectations hypothesis and to analyse whether the short-run changes $(y_{n-m,t+m} - y_{nt})$ in the long-term interest rate y_n are in line with the expectations theory:

$$(y_{n-m,t+m} - y_{nt}) = \alpha + \beta_n * (\frac{m}{n-m}) * (y_{nt} - y_{mt}) \tag{13}$$

Where $n = 120$ months is the maturity of the long-term bond and $m = 3$ months is the maturity of the short-term bond. At time t the long bond has a remaining maturity of m, at time $(t+m)$ at which the short bond matures it has a remaining maturity of $(n-m)$. Therefore, in order to adjust for the change in the remaining maturity, the term spread $(y_{nt} - y_{mt})$ has to be adjusted with the factor $(\frac{m}{n-m})$. For the expectations theory to hold β has to equal one. According to the expectations theory long rates should increase when the term

Table 5: Test of expectations hypothesis: regression of long-rate changes on term spread

	DE	ES	FR	IT	JP	NL	UK	US
α	-0.14	-0.32	-0.15	-0.32	-0.24	-0.13	-0.10	-0.17
t-stat	-4.2	-6.2	-4.2	-4.2	-4.3	-3.4	-2.9	-3.2
β_n	2.74	7.51	3.61	7.07	4.46	2.63	2.22	1.95
t-stat	3.5	5.7	4.3	3.5	2.9	3.0	2.8	1.8
Wald	0.03	0.00	0.00	0.00	0.03	0.07	0.13	0.39

estimation of equation 13, sample: 1990 - 2003

t-statistics are based on Newey-West autocorrelation consistent standard errors

Wald test: p-value for null-hypothesis that coefficient on term spread equals one

spread increases.

As table 5 points out we reject the hypothesis that the coefficient on the term spread is equal to one and, therefore, the expectations hypothesis for Germany, Spain, France, Italy and Japan at the five percent level. Although we reject the expectations hypothesis on the grounds of a purely statistical assessment, the estimation results show that the yield spread between ten-year bonds and three-month bills contains information for the change in the yield of the long-term bond during the life-time of the tree-month bill. For all countries with the exception of the US we find a significant and positive relationship between term spread changes and changes in the long rate even if the magnitude of the coefficient does not accord with the predictions of the expectations theory. The findings for the US do not come as a surprise because many studies reject the validity of the expectations theory for nearly all possible combinations of short and long US interest rates.

Shiller et al. [31] emphasise that the coefficient of regressions of the long-rate change on the term spread tends to be significantly different from one or even negative, particularly for combinations of relatively high n and low m. In contrast, the expectations theory seems to work better for low n and low m. One example for which the expectations theory seems to work quite well would be the term spread between two treasury bills of relatively short maturities.

With the application of present value models to the term structure, however, Campbell and Shiller [6] find that the expectations theory works fairly well even for high n and low m. In our case ($n = 120$ and $m = 3$) the coefficients of the term spread on long-rate changes display the correct sign but seem to be very far from one. We reach their same conclusion as we point out in table 5.

Table 6: Test of expectations hypothesis: regression of short rate changes on term spread

	DE	ES	FR	IT	JP	NL	UK	US
α	−2.61	−4.71	−3.46	−4.56	−3.40	−3.46	−3.28	−3.76
t-stat	−17.5	−20.7	−27.4	−12.0	−9.2	−24.4	−25.2	−18.5
β_m	1.07	0.94	0.87	0.70	1.62	0.91	1.02	1.36
t-stat	11.9	5.4	14.7	2.3	6.6	14.3	16.9	9.3
Wald	0.45	0.74	0.04	0.33	0.01	0.17	0.80	0.00

estimation of equation 14, sample: 1990 - 2003

t-statistics are based on Newey-West autocorrelation consistent standard errors

Wald test: p-value for null-hypothesis that coefficient on term spread equals one

Other work, for example Fama and Bliss [17], reports on the ability of the yield spread to forecast short rate changes over long horizons. Here the results seem to be more favourable for the expectations theory of the term structure. However, the predictive ability of the forecasts fades away when the forecasting horizon increases above nine months. Afterwards the predictive ability starts to improve again for horizons between two and five years as Fama [16] and Shiller et al. [31] report. The slope of the term structure can help to forecast interest rate changes over a few months but its predictive power decays rapidly over longer horizons.

$$\frac{\sum_{i=1}^{n-1} y_{m,t+i}}{n-1} - y_{mt} = \alpha + \beta_m * \frac{(n-m)}{n} * (y_{nt} - y_{mt}) \qquad (14)$$

As displayed by table 6, at the five per cent level we reject the expectations hypothesis only for France, Japan and the US.

The problem of unit roots in short rate interest rate changes is well known. Only for the UK short rate change we reject the hypothesis of a unit root at the 95 percent significance level. For the term spreads instead, we cannot reject the unit root for any of the countries at conventional levels of significance. Due to the low reliability of unit root tests and in line with the literature we essentially ignore this problem. Our contribution is not to test the validity of the expectations hypothesis but to analyse the theoretical and actual term spreads' reactions to monetary policy shocks.

Table 7: Summary statistics for standard tests of the expectations hypothesis

	DE	ES	FR	IT	JP	NL	UK	US
$\frac{\sigma(S^*)}{\sigma(S)}$	1.15	1.68	1.20	1.89	1.89	1.21	1.10	1.46
$\sigma(RP)$	0.78	1.68	0.93	1.86	1.18	0.78	0.99	1.15
ρ	0.87	0.70	0.86	0.45	0.73	0.87	0.86	0.79
β_n	2.74	7.51	3.61	7.07	4.45	2.63	2.22	1.95
β_m	1.07	0.94	0.87	0.7	1.62	0.91	1.02	1.36
β_{m-n}	1.00	1.18	1.03	0.86	1.38	1.04	0.94	1.15

Other standard tests of the expectations hypothesis Other often applied tests of the expectations hypothesis include the ratio of the standard deviations of the theoretical and the actual term spread and the standard deviation of the discrepancies between the two term spreads.

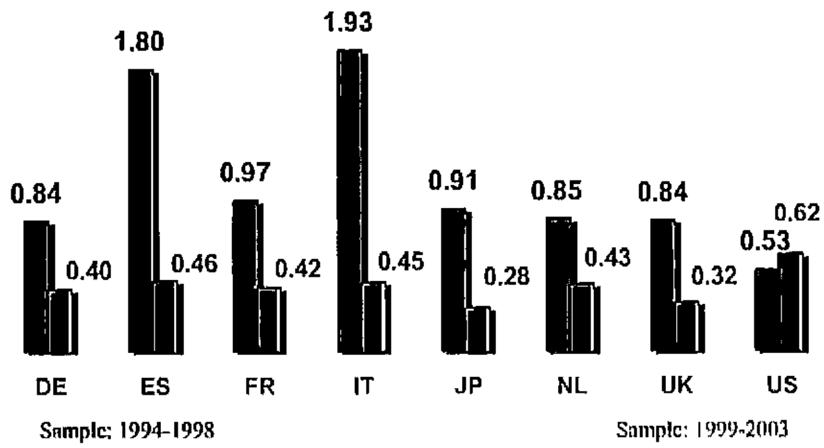
The ratio of standard deviations $\frac{\sigma(S^*)}{\sigma(S)}$ between the theoretical spread and the actual spread is much bigger than one for all countries in a sample going from 1990 through 2004. Thus, the variability of the term spread is too low to accord with the expectations model.

Table 7 summarises various test statistics for the validity of the expectations hypothesis. The sample correlation *rho* between actual and theoretical spread turns out to be very high for some countries but quite low for Italy. The same is true for the coefficient β of the regression of the theoretical spread on the term spread. This indicates that it is more problematic to apply the expectations hypothesis in Italy than let us say in the UK or in the Netherlands.

The sample standard deviation of discrepancies between the actual and the theoretical spread $\sigma(\varepsilon)$ quantifies the measurement error of the expectations theory. If the expectations theory were to hold, it should be small in relation to the standard deviation of the term spread.

The monthly standard deviations of the risk premia in table 7 are quite high but they tend to overestimate the variability of the risk premia since the introduction of the euro. This can be shown when the analysis of the variability of the risk premia is broken up in two subsamples. The standard deviations of the risk premium were about 110 bp between 1994 and 1998, but only around 40 bp afterwards. The variability of the risk premium of each country is displayed by figure 5. These risk premia are of considerable size when put into context of a standard deviation of the term spread of approximately 100 bp between 1994 and 1998 and about 40 bp afterwards.

Figure 5: Monthly variability of the risk premia before and after monetary unification



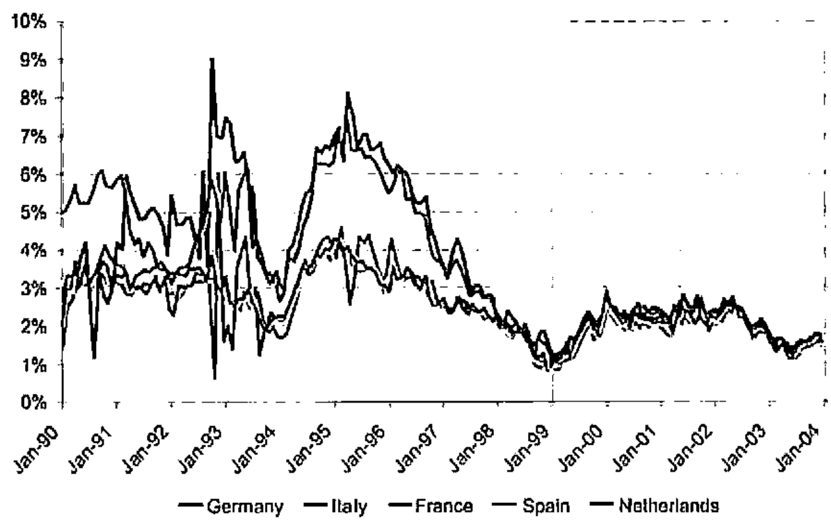
Bekaert and Rodrick (2000) [4] point out three different reasons for which the expectations hypothesis could be rejected. First, irrational investor behaviour might cause systematic forecast errors. Second, the presence of time-varying term premia means that tests of the expectations hypothesis are subject to omitted variable problems and estimated coefficients are biased. Third, tests may lead to false rejections due to their poor finite sample properties.

The statistical evidence in this section, overall, suggests the rejection of the expectations theory of the term structure for the eight countries studied. However, the expectations theory still yields useful information on the dynamics of interest rates in a more informal context.

4.3.2 The Collapse of the Risk Premium

The deviations of the ten-year interest rate from its expectations theory-based measure, according to the Campbell Shiller model in equation 1, can be interpreted as the risk premium inherent to ten-year interest rates over three-month interest rates. This risk premium contains the term premium of long bond yields over short bond yields in any specific country and a country-specific risk premium. Furthermore, the risk premium might contain other risk factors, the

Figure 6: EMU risk premium



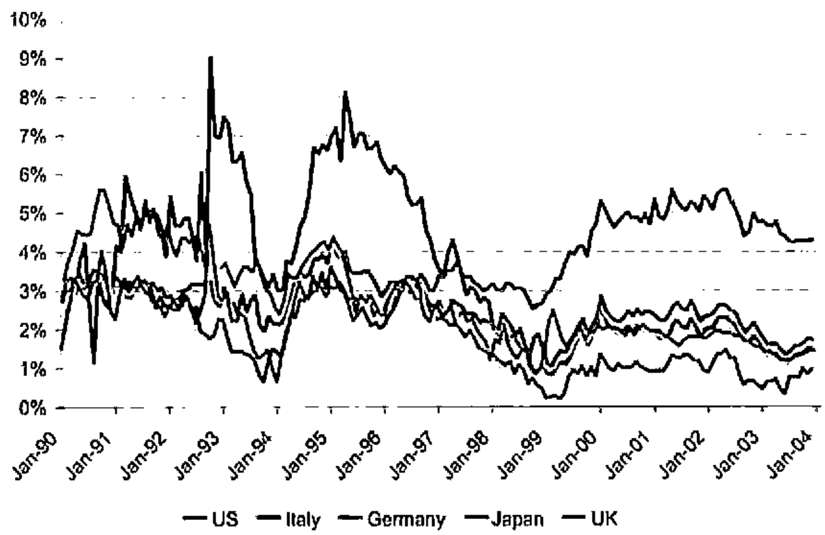
most prominent probably being the liquidity risk.

The risk premium between a short bond maturing after 3 months and a long bond maturing after 120 months, calculated using the weighted theoretical spread from equation 8, is depicted in figure 6.

The risk premia in EMU countries collapse in the second half of the 1990s on the road to monetary union, with only a few basis points of difference remaining at the end of our sample. This collapse in the risk premia has substantially contributed to the greater homogeneity perceived in the process of monetary policy transmission in the eurozone countries after 1999.

In the next section we will show that the reaction of theoretical, risk-adjusted, term spreads to monetary policy shocks has not fundamentally changed between the pre-EMU and the EMU period. However, the reactions of actual term spreads including the risk premia have changed fundamentally. The change that took place is attributable to the dynamics of the risk premium. A much lower and a much more uniform risk premium in the five eurozone countries of our study has contributed to increased homogeneity in the reaction to monetary policy shocks after 1999. For the homogeneity of the monetary policy transmission mechanism this seems to have played a much more important role than financial market structure or monetary policy signal quality effects.

Figure 7: International risk premium



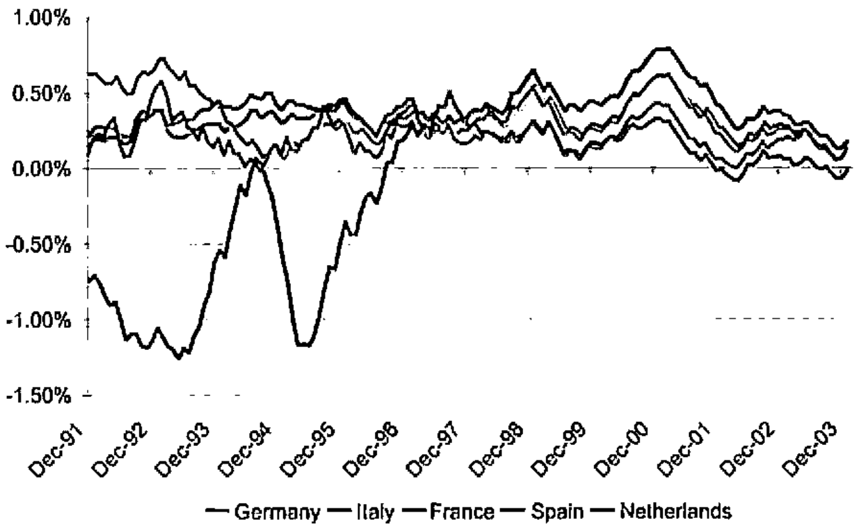
What has exactly happened to the risk premium in 1999? Durre et al. [12] test the hypothesis of a level shift in the constant risk premium in the German money market and their results suggest that the constant risk premium in the money market decreases for all maturities after the launch of the Euro in 1999.

For ten-year rates, however, the hypothesis of a constant risk premium has to be rejected and, more importantly, no decrease in the German risk premium can be observed.

The collapse of the risk premia of the five big euro area economies is in contrast to the considerable differences which remain in international risk premia, as shown by 7.

Another measure that has been widely used as a measure of the risk premium are Asset Swap Spreads measuring the premium paid by the banking system with respect to yields on government bonds. For asset swap spreads, too, there has been strong convergence during the build-up to monetary unification as displayed in figure 8 for the spread of ten-year swap rates over ten-year government bond rates from DataStream. Although there is no collapse, as the one observable for the risk premia, the correlation of monthly swap spreads makes a big jump in 1999. While between 1991 and 1998 the highest bilateral correlation

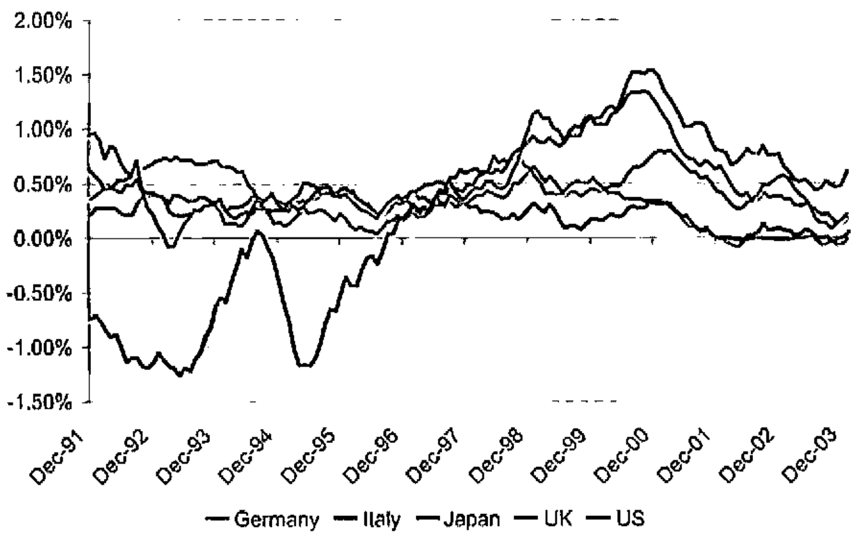
Figure 8: EMU Asset Swap Spreads (five month moving average)



pair is a correlation coefficient of 0.66 between the Netherlands and Germany, for the period from 1999 to 2003 the lowest observable correlation pair is a correlation coefficient of 0.81 between Germany and Spain. Nevertheless, differences in asset swap spreads continue to persist well into the monetary union. These differences represent heterogeneity in the riskiness of the banking systems and in the riskiness of government bonds. Between 1999 and 2004 we observe a persistent differential of the swap spreads of Germany and Italy which fluctuates in a band of 20 to 40 basis points. Apart from the different risk components contained in the three-month interbank rates of the two countries, the spread seems to derive from the higher risk premium inherent to Italian ten-year BTPs compared to ten-year German Bunds.

Clearly, the dynamics of the risk premia and of the asset swap spreads indicate a strong degree of convergence between the five EMU member countries during the period studied. In the next section we analyse whether the convergence of financial market asset prices implies as well a higher degree of homogeneity with respect to monetary policy shocks.

Figure 9: International Asset Swap Spreads (five month moving average)



5 Monetary Policy Transmission and the Risk Premium

5.1 Interest Rate Co-Movement

Among the five eurozone countries dealt with in our analysis, most correlations of monthly term spread changes adjusted for the risk premium displayed in table 8 are lower than correlations of actual term spread changes including the risk premium (as displayed before in section 3 in table 3). Roughly, this is true in the pre-EMU and in the EMU period and can be observed as well for the correlations with the UK which does not make part of the eurozone. Due to the high degree of economic integration of the UK with some eurozone countries, the risk premium of eurozone countries and the UK may contain a common 'eurozone' component. A common risk component inherent to the term spreads obviously increases the correlation between the term spread changes.

The time series dynamics in the correlations of risk-adjusted monthly term spread changes between the EMU and the pre-EMU period is very similar to the dynamics observed in the correlations of non-risk-adjusted term spread changes

Table 8: Correlation of monthly changes in risk-adjusted term spreads before and after monetary unification

	DE	ES	FR	IT	JP	NL	UK	US
DE	1	0.75	0.78	1.00	0.19	0.79	0.44	0.69
ES	0.15	1	0.96	0.75	0.15	0.51	0.27	0.51
FR	0.02	0.19	1	0.78	0.23	0.54	0.23	0.51
IT	0.25	0.40	−0.02	1	0.18	0.81	0.45	0.71
JP	0.17	−0.04	−0.20	−0.08	1	0.30	0.04	0.26
NL	0.63	0.31	0.06	0.41	0.10	1	0.32	0.59
UK	0.27	0.17	−0.03	0.36	0.08	0.40	1	0.54
US	−0.04	−0.07	−0.10	−0.15	0.20	−0.08	0.07	1

upper triangle: correl. of mthly. chg. in theoretical spread Jan'99-Jan'04
lower triangle: correl. of mthly. chg. in theoretical spread Jan'90-Dec'98
correlations between eurozone countries are reported in bold

in table 3. The increase in the correlation pairs was mainly attributable to the increase in the correlations of theoretical spreads of EMU with control group countries and between the three control group countries themselves[2]. Therefore, the convergence of monthly changes in theoretical spreads through time seems to be a phenomenon of global financial integration rather than a result of exclusive eurozone financial integration.

Another interesting piece of evidence is the correlation of monthly changes in risk-adjusted ten-year real rates. In section 3 we analyse ten-year real rates and find that the correlations between ten-year real rate changes have been increasing between eurozone countries themselves but they have been decreasing between eurozone and non-eurozone countries as displayed in table 2.

As displayed by table 9, by and large, correlations of risk-adjusted ten-year real interest rates increase between all countries in the period from 1990 to 2003 no matter whether a country makes part of the eurozone or not. The only exception are the Netherlands.

This is in contrast to not risk-adjusted ten-year real rate changes whose correlations only increase between eurozone countries themselves in the above-

[2]Between eurozone countries themselves correlations of changes in theoretical spreads actually increased by 3% less than did changes in the correlations of actual spreads in the EMU period with respect to the pre-EMU period. In contrast, the correlations in the changes of the theoretical spread between EMU countries and control group countries and between control group countries themselves increased 8% more than the correlations of the changes in the actual spread.

Table 9: Correlation of monthly changes in risk-adjusted ten-year real interest rates before and after monetary unification

	DE	ES	FR	IT	JP	NL	UK	US
DE	1	0.28	0.57	0.66	0.06	0.22	0.22	0.50
ES	0.04	1	0.42	0.29	0.18	−0.10	−0.01	0.27
FR	0.12	−0.01	1	0.35	0.10	−0.12	0.18	0.27
IT	−0.20	0.28	−0.25	1	0.30	0.18	0.19	0.57
JP	0.17	−0.11	−0.17	−0.05	1	0.15	0.05	0.15
NL	0.30	0.25	−0.02	0.20	0.16	1	−0.10	0.33
UK	−0.12	0.01	−0.03	0.18	−0.06	0.25	1	0.29
US	0.10	−0.01	−0.05	−0.12	−0.03	0.06	0.00	1

upper triangle: correl. of mthly. chg. in risk-adj. 10y real rates Jan'99-Jan'04
lower triangle: correl. of mthly. chg. in risk-adj. 10y real rates Jan'90-Dec'98
correlations between eurozone comtries are reported in bold

mentioned period but not between eurozone countries and control group countries, where they actually decrease as displayed by table 2. Clearly, looking at risk-adjusted ten-year real rates gives a much more homogeneous picture of interest rate dynamics between euro and control group countries. The abstraction from the risk premium leads to an increased degree of homogeneity in agents' interest rate and inflation expectations.

This subsection raises the point that eurozone countries and the UK could have a common risk premium component. Furthermore, data on term spreads and on ten-year real interest rates suggest increased financial integration between EMU countries themselves but as well between EMU countries and control group countries. It would seem that higher correlations in rates and spreads are not exclusively due to EMU because they include non-EMU countries, too. We would expect that 'risk-adjusted' prices of financial assets in euro area countries will not display different reactions to monetary policy shocks in the period between 1999 and 2004 compared to their reactions in the period from 1990 to 1998.

5.2 Monetary Policy Shocks

We analyse the reaction of the term spreads of ten-year government bond rates over three-month money market rates to monetary policy shocks before and during EMU in order to show that the reactions of the theoretical spreads have not changed significantly, while the reactions of the actual spreads have changed fundamentally. The difference between the two spreads being the risk premium, it seems to be the risk premium which has caused the change in the responses of the term spreads to monetary policy shocks after 1999.

We estimate a vector-autoregressive (VAR) model of the form

$$y_t = A_1 y_{t-1} + ... + A_p y_{t-p} + Bx_t + e_t \tag{15}$$

where y_t is the vector of the endogenous varibles

$$y_t = \begin{pmatrix} \Delta i^*_{t,t-1} \\ \Delta i_{t,t-1} \\ TS_t \end{pmatrix} \tag{16}$$

and where x_t is a vector of exogenous variables and $A_1, ..., A_p$ and B are the matrices of the coefficients to be estimated and e_t is a vector of innovations that may be contemporaneously correlated but are uncorrelated with all the right-hand side variables.

We estimate the VAR in the first difference of the German short-term interest rate $\Delta i^*_{t,t-1}$, where appropriate (for France, Italy, Spain, the Netherlands and the United Kingdom), the first difference in the domestic short-term interest rate $\Delta i_{t,t-1}$, and the domestic yield spread of ten-year government bonds over three-month money market rates TS_t. We add a constant as the only exogenous variable and choose two lags. For Germany, Japan and the US the VAR will be just a bivariate VAR in the first difference of the domestic three-month money market rate and in the yield spread.

Apparently, we are not using the typical monetary policy VAR which probably would include inflation and output, too, when analysing monetary policy shocks. This is because we stick to the methodology of Campbell and Shiller [6] used for forecasting the short-term rate changes. This methodology is used to calculate the risk premia and, therefore, we stick to this methodology when analysing monetary policy shocks to term spreads and risk-adjusted term spreads.

The Cholesky identification scheme used to identify the single shocks of each equation has the ordering TS_t, $\Delta i_{t,t-1}$, $\Delta i^*_{t,t-1}$. The foreign short-term interest rate is ordered last because there should be no feedback from the domestic short-term interest rate or the domestic term spread to the foreign short-term interest rate. The domestic short-term rate is ordered second because the term spread can be expected to react to domestic monetary policy shocks. Finally, the term spread is ordered first as it is shocked by both the other variables in the system. The spread between $10y$ bonds and $3m$ treasury bills should react negatively to changes in monetary policy rates. This would be in line with monetary policy rates raising rates at the short-end but having no immediate effect on the long-end of the yield curve. Monetary policy's effects on the long-end work via inflationary expectations which do not only depend on monetary policy but as well on many other factors like the wage/price setting process, general labour market conditions, commodity prices, productivity and exchange rate dynamics, to name just a few of them.

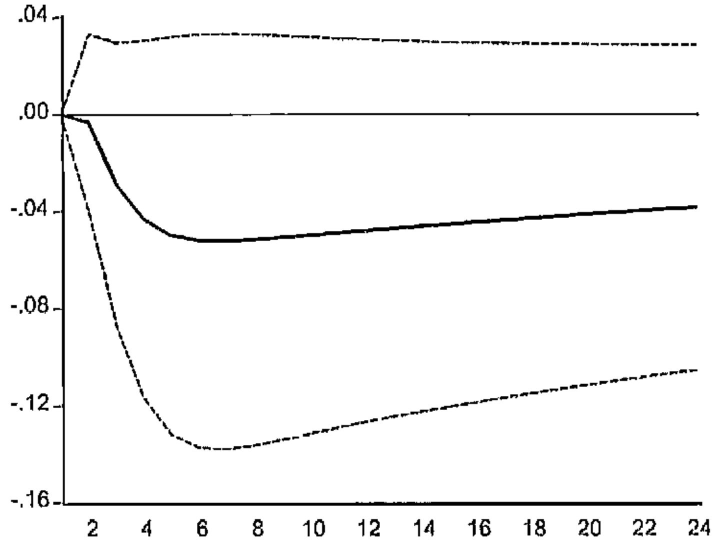
Actual spreads, in general, react stronger to the policy rate shocks than theoretical spreads do, as displayed in figures 10 to 29. However, the difference in the magnitude of the reactions is more marked before the introduction of the euro and then decreases after the introduction of the single currency. An exception to this is Germany, and to some degree the US, where reactions of theoretical spreads to monetary policy shocks are slightly stronger than reactions of actual spreads.

Importantly, after the introduction of the euro actual spreads in all countries start to react in a consistent and statistically significant way to monetary policy innovations. This is true for risk-adjusted spreads, too, but it is particularly true for actual spreads as figures 10 to 29 point out. Actual spreads of the eurozone countries converge up to a very high degree after the euro introduction because the elimination of the exchange rate risk lets collapse the risk premia inherent to the single member states long-bond rates. Therefore, the risk-adjusted spreads react more uniformly to policy shocks after the advent of the euro.

For the case of Germany the actual spread seems to react stronger to policy signals in the EMU period compared to the pre-EMU period (see figures 10 and 11). This seems to be in line with an increased speed of monetary transmission and pass-through from policy rates to market rates after the euro introduction. Looking at the risk-adjusted spread instead, the reaction to monetary policy shocks is already quite strong in the pre-EMU period. Essentially, the response of the theoretical spread does hardly change between the pre-EMU and the EMU period. If anything, the persistence of the shocks to the theoretical spread seems to have been slightly higher in the pre-EMU period.

The Italian term spread hardly reacts to the innovations from the policy rate of the Bank of Italy but shows instead a stronger reaction to the policy shocks of the Bundesbank in the period from 1990 to 1998. Figure 14 displays

Figure 10: Response of actual German term spread to one S.D. shock in German policy rate (1990-1998)



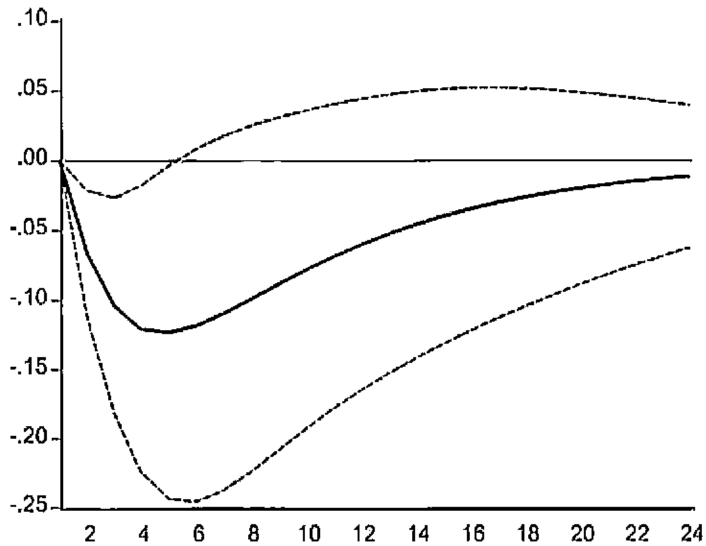
the reactions of the Italian actual term spread to Italian and German monetary policy shocks.

The term spread during EMU reacts in the expected manner to the ECB policy signal as displayed by figure 15. There has been a big change in the reaction of the Italian term spread to monetary policy surprises after the introduction of the Euro.

A close look a the reactions of the theoretical spread shows that in contrast to the actual spread, essentially, there was only a change in the magnitude of the reactions after 1999. The Italian theoretical term spread showed the expected negative reaction to the Italian monetary policy rate already in the period between 1990 and 1999, as displayed in figure 16. During the euro regime the theoretical Italian term spread in figure 17 shows a very similar reaction to monetary policy. This suggests that in 1999 no change took place in the structural relationship between monetary policy rates and the theoretical term structure, as shown by comparing figures 16 and 17.

The very different behaviour of the theoretical and the actual term spread is due to collapse of the risk premia in 1999. The reaction of the actual term

Figure 11: Response of actual German term spread to one S.D. shock in ECB policy rate (1999-2004)

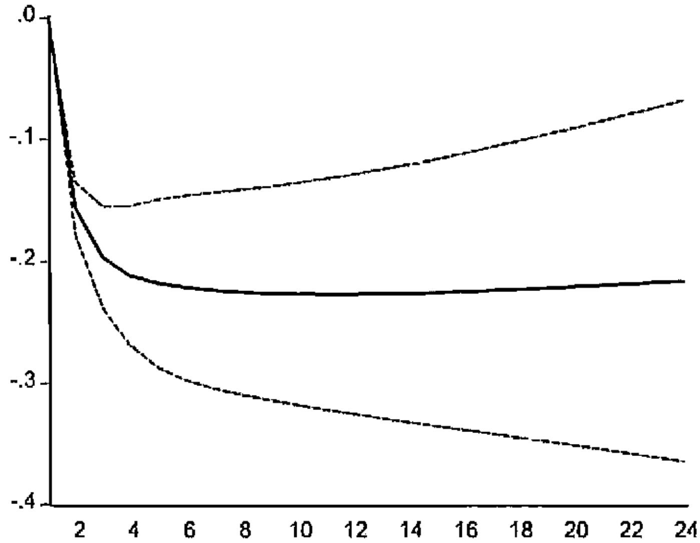


spread has changed in reaction to the decline in the risk premium due to the elimination of the exchange rate risk.

Spain confirms the abnormal reaction of the actual term spread to monetary policy signals of the Italian case in the sense that the Spanish term spread shows the wrong, positive, reaction to the Spanish policy shock as displayed by figure 18. A positive reaction of the term spread means that long rates increase more than short rates in reaction to the monetary policy shock. This might be testimony of the markets pessimism towards the conduct of Spanish monetary policy. The fact that the long rate increases more than the short rate in reaction to the monetary shock might be interpreted as agents' expectations that long rates have to increase in order to incorporate a higher inflation premium. In contrast, figure 18 suggests that a German monetary policy shock produces the correct, negative, reaction of the Spanish term spread. Thus, a more restrictive stance to monetary policy in Germany forces a tightening of market liquidity in Spain, too. Figure 19 shows, instead, that the reaction of the Spanish spread in the EMU period displays the intuitively correct, negative, sign.

The impulse responses of the theoretical Spanish term spread in figures 20

Figure 12: Response of theoretical German term spread to one S.D. shock in German policy rate (1990-1998)



and 21 underline the fact that the dynamics of the theoretical spread did not change qualitatively between our two sub-periods pre-EMU (1990-1998) and EMU. The positive reaction of the Spanish actual term spread to the Spanish monetary policy shock is ascribable to the risk premium inherent to the long end of the Spanish yield curve. The risk premium component in the actual term spread reacts positively to a Spanish interest rate hike because markets could interpret it as a defence of the Spanish exchange rate by the hands of the Spanish central bank. This component breaks away only with monetary unification in 1999. Clearly, the Spanish situation is very similar to the Italian one analysed above and can be compared to the French and the Dutch experience (figures 22 to 29).

Figure 13: Response of theoretical German term spread to one S.D. shock in ECB policy rate (1999-2004)

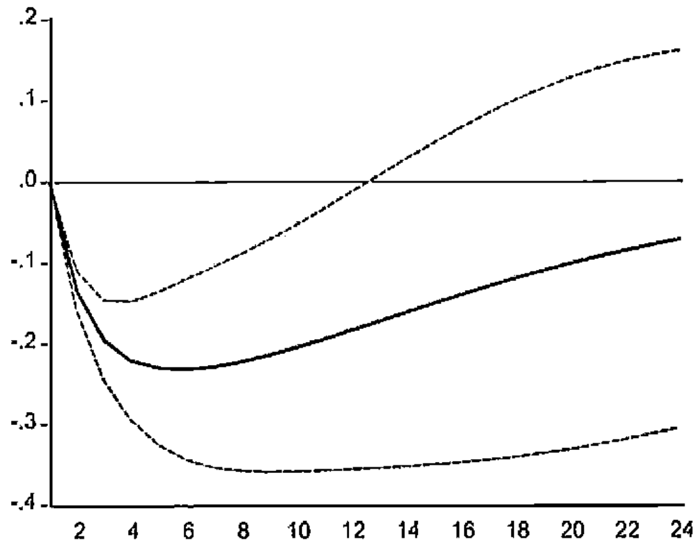


Figure 14: Response of actual Italian term spread to one S.D. shocks in Italian and in German policy rate (1990-1998)

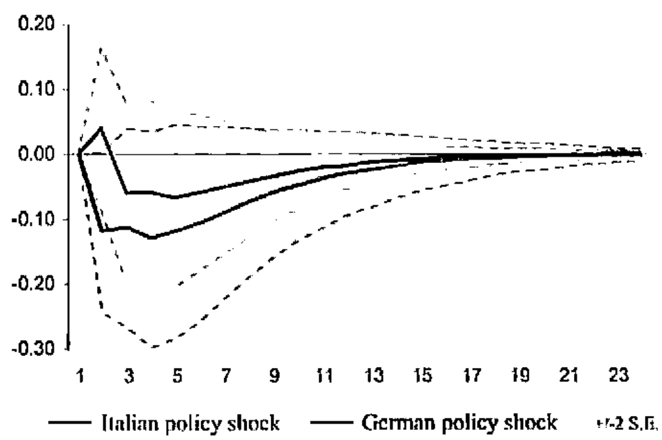


Figure 15: Response of actual Italian term spread to one S.D. shock in ECB policy rate (1999-2004)

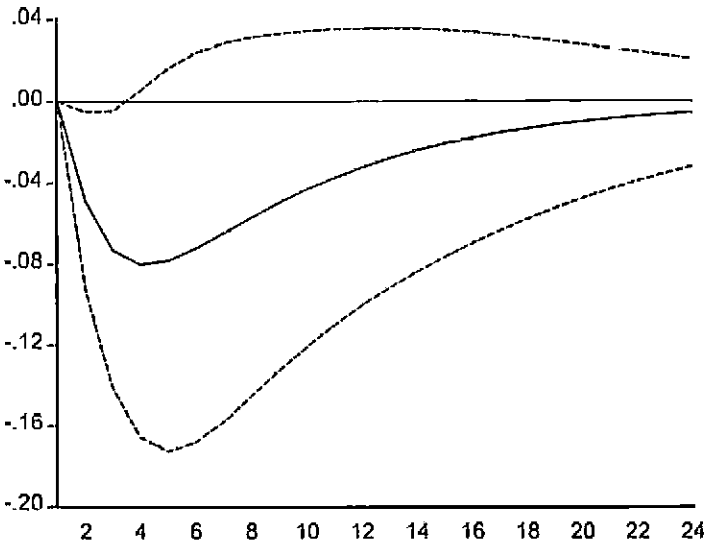


Figure 16: Response of theoretical Italian term spread to one S.D. shock in Italian and in German policy rate (1990-1998)

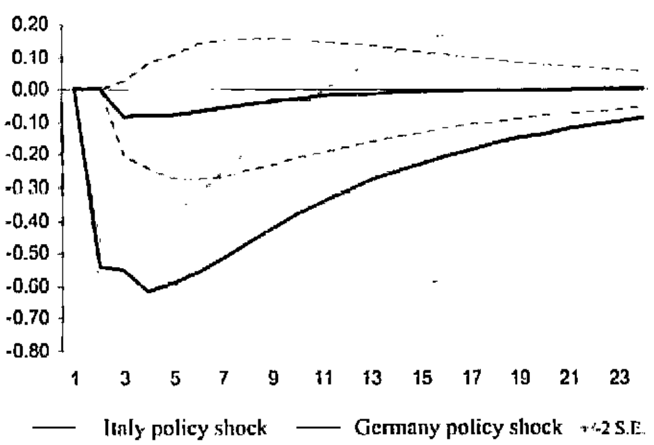


Figure 17: Response of theoretical Italian term spread to one S.D. shock in ECB policy rate (1999-2004)

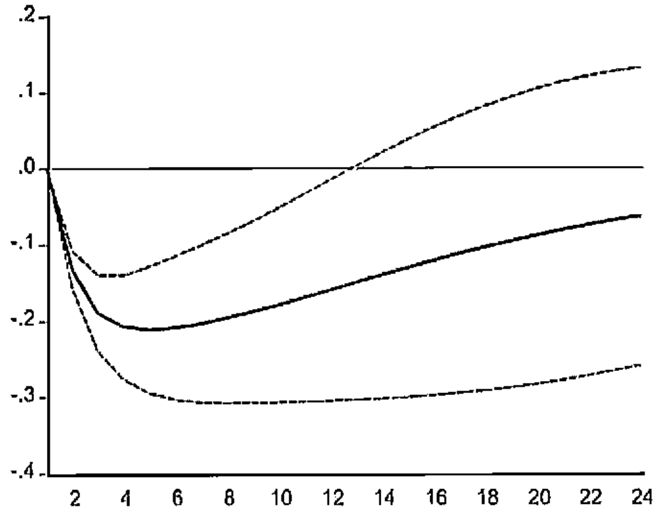


Figure 18: Response of actual Spanish term spread to one S.D. shocks in Spanish and in German policy rate (1990-1998)

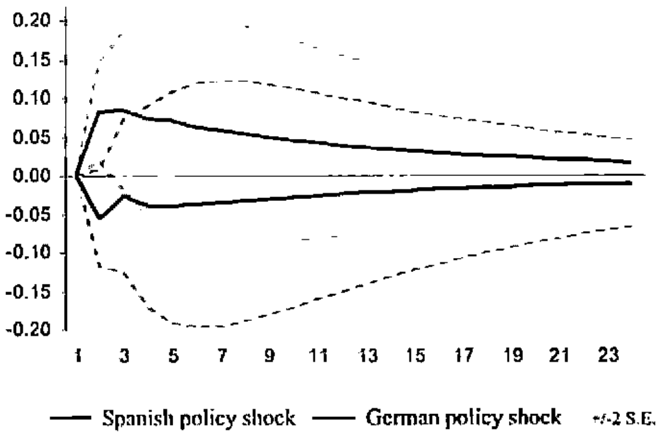


Figure 19: Response of actual Spanish term spread to one S.D. shock in ECB policy rate (1999-2004)

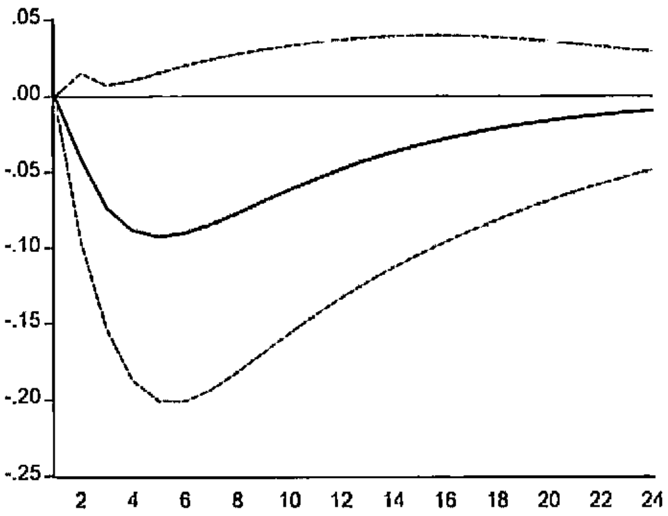


Figure 20: Response of theoretical Spanish term spread to one S.D. shock in Spanish and in German policy rate (1990-1998)

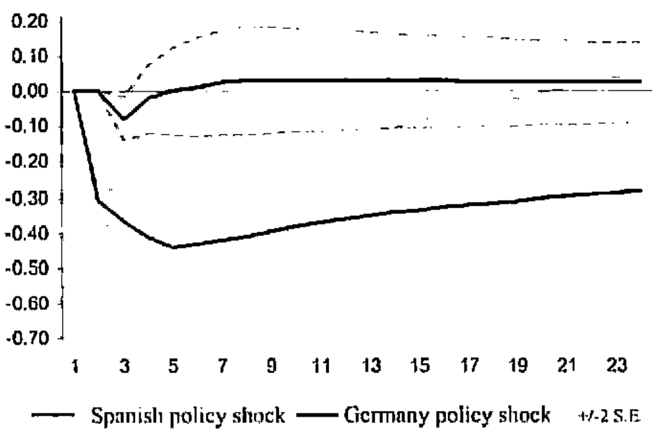


Figure 21: Response of theoretical Spanish term spread to one S.D. shock in ECB policy rate (1999-2004)

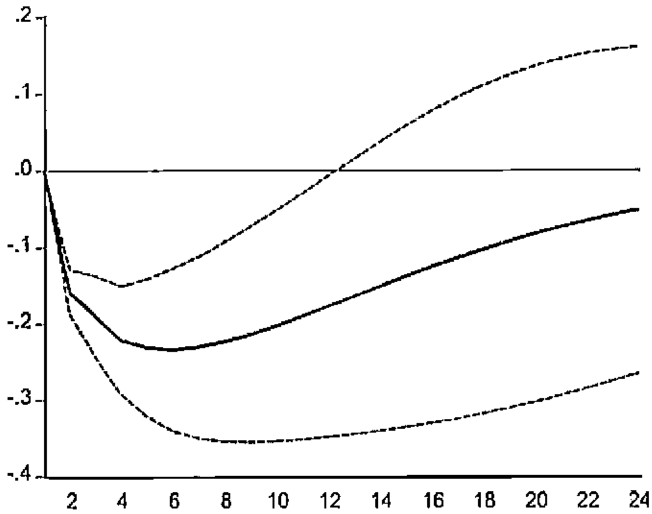


Figure 22: Response of actual French term spread to one S.D. shock in French and in German policy rate (1990-1998)

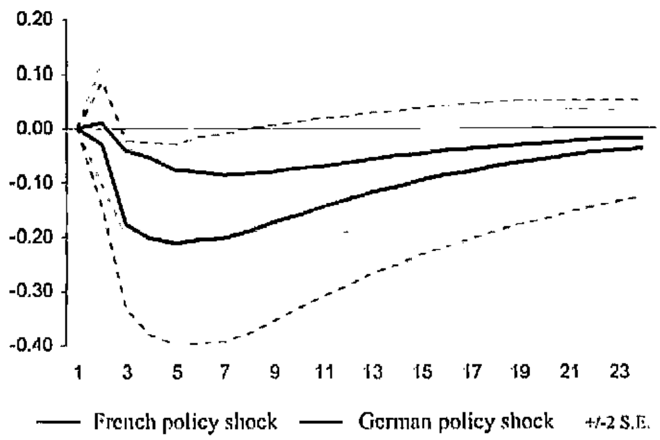


Figure 23: Response of actual French term spread to one S.D. shock in ECB policy rate (1999-2004)

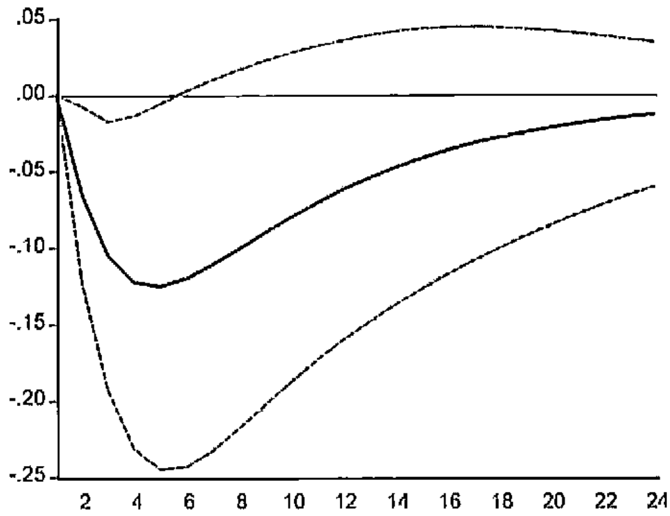


Figure 24: Response of theoretical French term spread to one S.D. shock in French and in German policy rate (1990-1998)

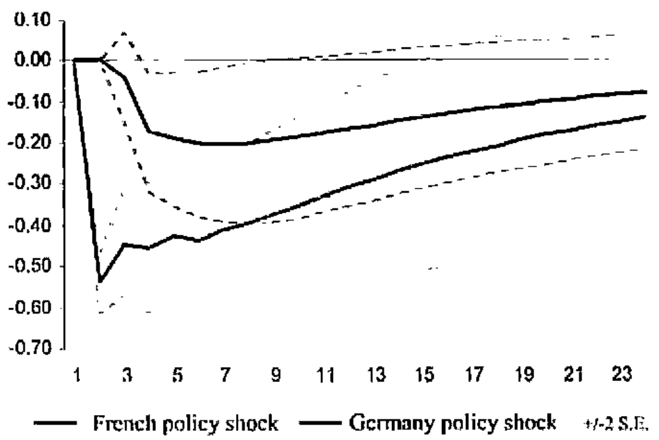


Figure 25: Response of theoretical French term spread to one S.D. shock in ECB policy rate (1999-2004)

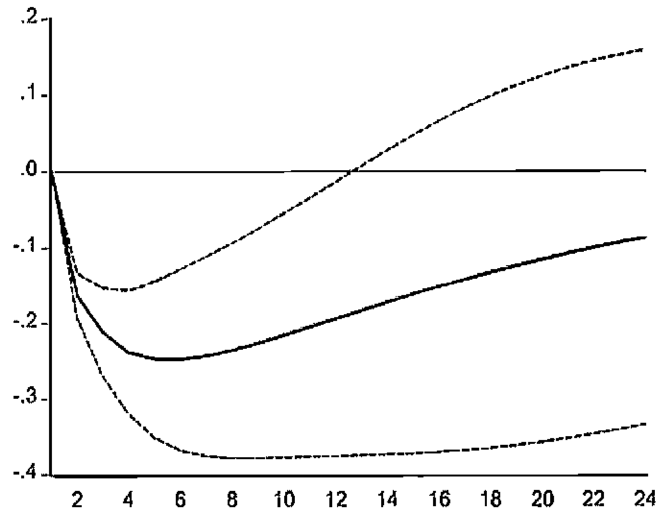
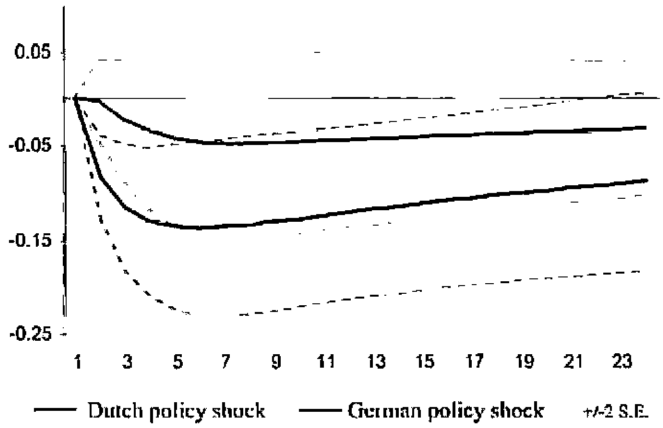


Figure 26: Response of actual Dutch term spread to one S.D. shock in Dutch and in German policy rate (1990-1998)



In all four EMU countries analysed the actual term spread shows a stronger reaction to German monetary policy shocks than to domestic monetary policy shocks in the period between 1990 and 1998. In some countries the domestic shock initially even causes the national term spread to increase, instead of tightening it as it would be expected from in reaction to positive monetary innovations. For all countries the somehow strange reaction of the actual term spread to monetary policy shocks then gives way to more conventional reactions after 1999.

The reactions of the theoretical spread, on the other hand, seem to be standard in all countries and are qualitatively the same for the pre-EMU and the EMU period. Quantitatively they are very similar for the two sub-periods, too.

Figure 30 points out how the risk premium inherent to Italian ten-year government bonds reacts to monetary policy shocks before EMU and during the EMU-period. It shows that there was a very strong reaction of the risk premium before EMU, probably reflecting agents' inflationary expectations. The reaction is much milder during EMU. The strong jump in the reaction of the risk premium to monetary policy shocks in the before-EMU period stands behind the seemingly surprising result from figure 14 that monetary policy shocks have initially a positive effect on the actual Italian term spread. The reaction

Figure 27: Response of actual Dutch term spread to one S.D. shock in ECB policy rate (1999-2004)

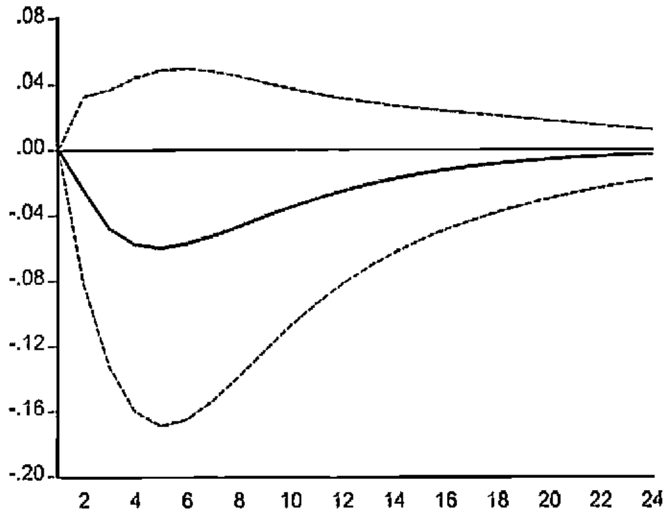


Figure 28: Response of theoretical Dutch term spread to one S.D. shock in Dutch and in German policy rate (1990-1998)

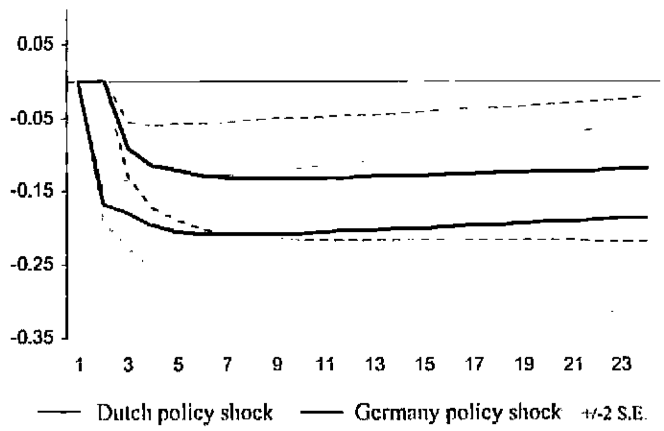
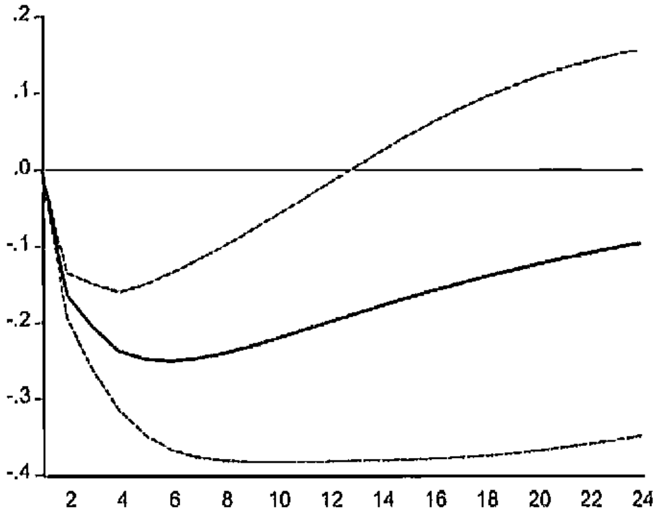


Figure 29: Response of theoretical Dutch term spread to one S.D. shock in ECB policy rate (1999-2004)

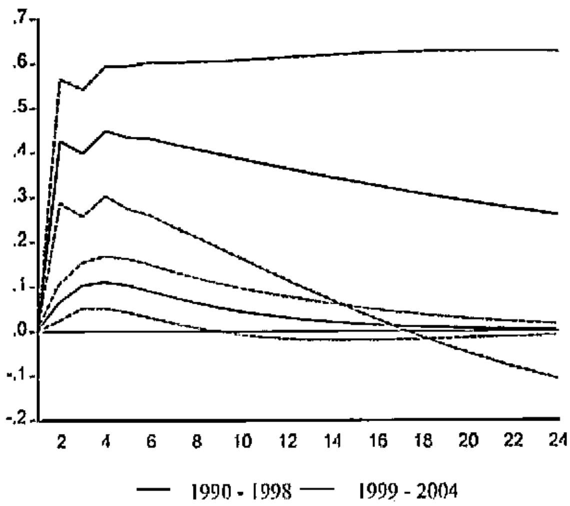


of the risk premium wholly explains the different reactions of the actual term spreads to monetary policy shocks before and during EMU. This does not pertain exclusively to monetary policy shocks in Italy but is representative for all five eurozone countries covered by our study. The collapse of the risk premium alone is capable of explaining a higher degree of homogeneity in monetary policy transmission since the advent of the euro.

The considerable asymmetry of the effect of monetary policy shocks between countries pointed out above fades away with the advent of EMU. Now, the risk premia component of yields in the five eurozone countries display roughly identical dynamics and, therefore, the reaction of actual spreads to monetary policy shocks become qualitatively and quantitatively very similar. The same happens, to monetary policy shocks in the UK. Even if the UK makes not part of the euro area its risk premium during the EMU-period displays a much higher correlation with the risk premia of the eurozone countries than it did in the pre-EMU period.

Our results clearly show that, higher homogeneity in monetary policy shocks and monetary policy transmission is explained up to a high degree by the collapse in the risk premium even without considering evolutions in financial mar-

Figure 30: Reactions of the term premium inherent to ten-year Italian government bond yields to monetary policy shocks



ket structure, which require extensive time periods to evolve, or changes in the monetary policy signal.

All eurozone countries show qualitatively similar reactions of their risk-adjusted term spreads to monetary policy shocks. Reactions of actual spreads to monetary policy shocks, however, do vary a lot between the pre-EMU and the EMU period. The strong fall in the risk premia causes 'more normal' reactions of the yield curve to monetary policy shocks in Italy, France, the Netherlands and Spain. Markets in these countries used to react with increases in long-rates to monetary policy shocks by central banks. Increases in policy rates where often interpreted as manoeuvres to defend the exchange rate and markets reacted by requiring higher yields on long-term debt.

6 Conclusions

There seems to be evidence for increased homogeneity in the monetary policy transmission process of the euro area countries in the aftermath of monetary unification in 1999. Concerning the interest rate channel of monetary transmission, three-month real interest rates, ten-year real bond yields and the term spread of ten-year government bonds over three-month money market rates all have become increasingly correlated during the EMU period compared to the period from 1990 to 1998. With the exception of the ten-year real bond yields, this has, however, been the case for the correlations between eurozone countries and countries of a control group, made up of Japan, the US and the UK, too. Increased co-movement in interest rates, asset prices and probably in banking behaviour has taken place as well between the eurozone economies and the three other big economies not belonging to the euro area and, therefore, deeper financial integration in the eurozone as represented by interest rate dynamics is no sufficient proof of increased homogeneity in the eurozone monetary policy transmission mechanism.

Nevertheless, monetary transmission most likely has become more homogeneous in the wake of EMU. However, the signal-quality effects and market structure effects cited in the literature are probably not, on their own, capable of explaining increased homogeneity, neither seem they to be necessary in order to explain increased homogeneity. An important contribution clearly comes from the collapse of the risk premia in the eurozone countries during the second half of the 1990s. The collapse and the convergence of the risk premia of the euro area countries' ten-year government bonds up to 1999 has important implications for interest rate co-movement and asset pricing. Euro area agents now face a common yield curve when taking intertemporal decisions. Consequently, the collapse in the risk premia fundamentally changes the reactions of the term spreads to monetary policy shocks in the eurozone countries and suggests a higher degree of homogeneity in the monetary transmission mechanism.

A higher degree of homogeneity in the monetary transmission process leads to more symmetric propagation of shocks across eurozone countries after 1999. Looking at the reaction of risk-adjusted term spreads to monetary policy shocks we show that the reactions of theory-based measures of the term spread to shocks are already very similar before 1999 across eurozone countries. The reactions of the actual spreads to monetary policy shocks, instead, change considerably during the EMU period compared to the pre-EMU period. This is because policy shocks during the pre-EMU period lead to jumps in risk premia. In contrast, this is not the case anymore since the introduction of the euro. Thus, removing the risk premium component from term spreads, no major change in the reaction of term spreads to monetary policy shocks can be observed when the EMU period is compared to the pre-EMU period. This underpins our hypothesis that the increased interest rate co-movement and the, probably, increased homogeneity in monetary policy transmission can be accounted for by the collapse and the convergence of the risk premia rather than by a higher perceived quality of the policy signal or a market structure effect due to increased financial market integration.

Future research in this field should take care of the fact that studies comparing the evolution of the monetary policy transmission mechanism in euro area countries are challenging because they deal as well with the formation of agents' expectations across different monetary policy regimes.

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Appendix A: The Banking Channel of Monetary Policy Transmission

Banks and financial institutions generally play a crucial role in financial intermediation and, given the importance of bank lending as a source of finance in most eurozone countries, particularly so in the euro area. In contrast to structural econometric models which identify the transmission channels and their importance on the macroeconomic level, the behaviour of banks can supply important information on the credit channel on the microeconomic level. The importance of the bank-lending channel depends very much on the financial structure and on the financial stability of an economy. A priori it is not clear whether the bank-lending channel has important implications for monetary transmission in all eurozone countries. However, it might have important implications at least in some eurozone countries.

In this respect, some of the most important questions are: How do banks react to changes in policy rates and how do they pass on these changes to their customers and to other banks via interbank, deposit, lending, mortgage rates etc.? Can a breakpoint be found in bank behaviour that coincides with the introduction of the single currency in 1999 and did banks change their reaction to monetary policy because of EMU? What does this tell us about the homogeneity of the monetary transmission process in EMU? These are only some of the most crucial questions about the evolution of monetary policy transmission in the context of monetary unification.

Changes in the pass-through behaviour of banks might tell us whether monetary transmission via bank rates has changed over time. Concerning the direction of the change we are interested in finding out whether a more powerful or more rapid pass-through and transmission can be observed after the introduction of the single currency.

With the advent of EMU the nature of monetary transmission could have changed mainly due to two factors pointed out by Angeloni and Ehrmann [1].

Namely, these are the monetary policy signal quality effect and the financial market structure effect.

Firstly, the foundation of the ECB could have changed the quality of monetary policy signals or at least the perceived quality of these signals. For many countries monetary policy has become less volatile following its transfer from national central banks to the ECB. Therefore, the reactions of banks and financial institutions to a policy signal, now perceived as being of higher quality, are likely to have changed.

Secondly, the integration of the EU banking market has been further boosted by monetary union and could have affected the market structure of the banking industry in EMU countries. A higher degree of financial market integration is supposed to go hand in hand with less asymmetric reactions to shocks and, thus, increased homogeneity in the monetary transmission process. The market structure effect might be more sluggish to react than the signal-quality effect. Nevertheless, Angeloni and Ehrmann [1] point out that the duration of financial contracts is already converging in euro area countries, indicating a vivid market structure effect. Duration of financial contracts is falling in Germany and in the Netherlands, where, historically, it has been very long and is increasing particularly in countries like Italy and Spain, where it used to be very short.

Combined, the signal-quality effect and the market-structure effect could explain changes in the monetary transmission mechanism in EMU countries before and after monetary union. However, due to still scarce data on EMU it is a very tricky task to identify the market-structure effect and the signal-quality effect separately.

In order to get a proxy for, how powerful or how rapid transmission takes place we calculate the impact effects of money market rates for a selection of interest rates. The rates being used are lending rates, deposit rates and interbank rates which are each regressed on money market rates and on a constant.[3] Money market rates can be regarded as close substitutes of official rates as can be seen by figures 31, 33 and 32 which display three-month money market rates and official policy rates for Germany, the UK and Japan. The averages of coefficients on the money market rates for the five biggest eurozone economies and for three countries of a control group are reported in table 10.

A broad strand of emprirical literature on bank behaviour suggests that simple specifications are reasonably successful in capturing the dynamics of

[3] As pointed out by Angeloni and Ehrmann (2003), there should be no problems with omitted variables or endogeneity bias here.

Table 10: Average impact effects of money market rates on various banking rates in pre-EMU and EMU period

	1990 - 1998				1999 - 2003			
	coef	t-stat	const	t-stat	coef	t-stat	const	t-stat
DE	0.648	15.6	−0.002	0.2	0.442	5.9	−0.003	0.1
ES	0.109	3.9	−0.036	−1.1	0.359	2.3	−0.011	−0.5
FR	0.208	5.4	−0.029	−1.7	0.489	12.3	−0.006	−1.0
IT	0.427	21.9	−0.022	−0.9	0.558	22.5	−0.025	−1.7
JP	0.345	6.8	−0.025	2.3	0.224	2.3	−0.006	−1.7
NL	0.466	6.3	−0.012	−0.7	0.477	1.8	−0.032	−1.2
UK	0.607	19.4	−0.008	−0.4	0.657	8.0	−0.025	−1.2
US	0.529	14.6	−0.009	−0.5	0.602	11.7	−0.007	−0.7

bank rates as these are set administratively based on signals from the central banks.

Figures 31 to 32 suggest that in Germany but as well in Japan the transmission to the money market was much slower than in the UK for a good part of our sample. For Germany the picture changed only with the introduction of the euro. Since 1999 the money market rate adjusts much faster to changes in the policy rate and tracks the policy rate very closely. In the case of Japan the Japanese money market rate tracks the policy rate only closely since the second half of the 1990s when the Bank of Japan entered into a period of extremely low monetary policy rates to fight off deflation.

Concerning the EMU countries, the impact effects of the money market rate on a selection of market interest rates decreases between the pre-EMU and the EMU period only for Germany. On the other hand, the impact effects increase for Italy, Spain, France and the Netherlands. The standard deviation of impact effects between the five EMU countries decreases between the two sub-periods, but so it does for the control group, too. In the control group the impact effect increases slightly for the US and for the UK but decreases for Japan in the post-EMU period compared to the pre-EMU period.

One possible factor which could have influenced the impact coefficients in EMU in such a divergent way is the signal-quality effect due to the transfer of monetary policy from the national central banks to the ECB in 1999. In contrast to the above-mentioned market-structure effect, the signal-quality effect can be expected to have immediate impact on monetary policy transmission once a

Figure 31: Deutsche Bundesbank/ECB policy rate and German three-month money market rate

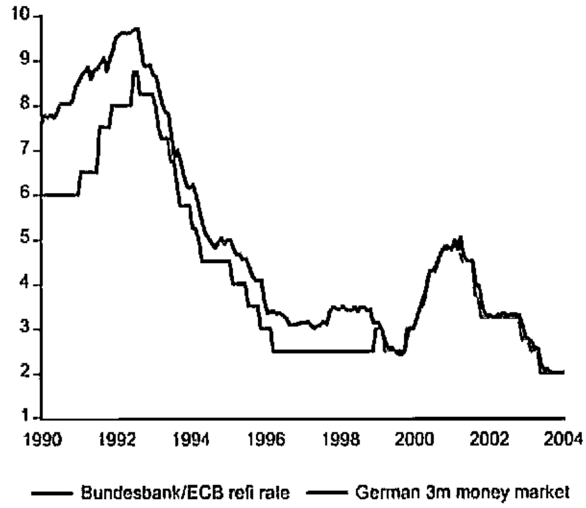


Figure 32: Bank of Japan policy rate and Japanese three-month money market rate

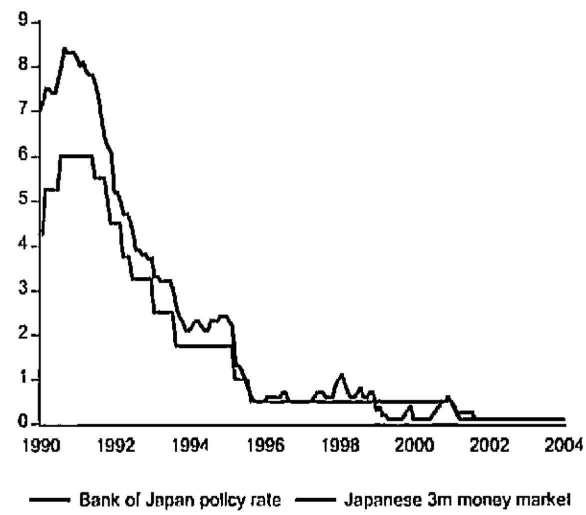
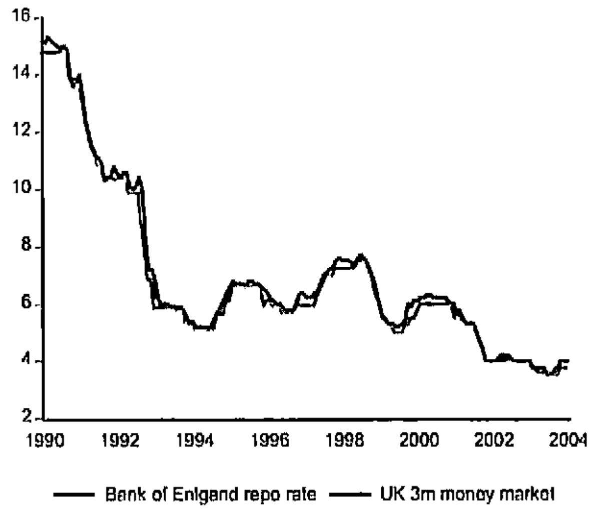


Figure 33: Bank of England policy rate and UK three-month money market rate



central bank changes operation procedures or once the institutional framework in which a central bank operates changes, like it did in the case of EMU.

In some countries the response of commercial banks to policy signals of any given size has increased, while in others it has decreased. If signal quality played a role in the evolution of commercial banks' reactions to policy rates this could imply that banks in Germany perceived that signal quality has suffered from the changeover of monetary policy from national central banks to the ECB in 1999, while banks in Spain, France and Italy perceived that signal quality has improved. The perception of the quality of a signal by agents depends on the observed variance of a monetary policy signal. When a signal contains too much noise, the response to the signal by agents will be reduced.

Even if the market-structure effect needs time to work itself through the financial system it might have already contributed to a more homogeneous transmission process in EMU. An important component of market-structure concerns the duration of financial contracts. Duration of financial contracts seems to have decreased in countries which used to have a very stable monetary policy like Germany and the Netherlands and seems to have increased in countries which were suffering from a lack of credibility in monetary policy like Italy and Spain.

The banking channel of monetary policy transmission seems to indicate that commercial bank behaviour has contributed to a higher degree of homogeneity in the transmission process. The impact effects of official policy rates on bank rates have increased in countries with historically low monetary policy credibility and decreased in countries with historically high credibility yielding a more uniform picture of commercial banks' interest rate setting process during EMU than during the pre-EMU period.

An interesting question would be to analyse whether expectations and perceptions about risk connected to the term structure might have caused levels of uncertainty about monetary policy and future short and long-term rates which are capable of explaining the sluggish reactions of the banking sector to monetary policy and, thus, the low impact effects of changes in money market rates between 1990 and 1998? Giving an answer to this question is beyond the scope of this paper, however.

Instead, we analysed the interest rate channel of monetary policy transmission and the implications the collapse the country risk premia had for the degree of homogeneity of transmission across euro area countries. While the development of the bank lending channel has contributed to a higher degree of homogeneity in the monetary transmission mechanism across countries we have shown that the financial markets' risk premium-adjusted reactions to monetary policy shocks, in contrast, did not change between the pre-EMU and the EMU period.

Does the Yield Spread Predict Output Growth in the CEE Transition Economies?

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December 2005

Abstract

We investigate the predictive content of the yield spread for future output growth in the CEE transition economies. Looking at two different measures of the yield spread - the yield spread of five-year government bonds over three-month treasury bills on the one hand, and the yield spread of twelve-month interbank rates over one-month interbank rates on the other hand - we decompose both measures of the yield spread in an expectations-related and a term premium component and find that innovations to both of the two components indicate future economic expansions in a model including the government spread, whereas they indicate future economic contractions in a model including the interbank spread. This seems to be true on the single country level as well as on the basis of a panel estimation approach.

Keywords : Yield Spread, Output Growth, Taylor rule, CEE Economies
JEL classification : E43, E52

1 Introduction

Historically, there has been a keen interest in analysing the term spread as one of the most significant leading indicators of future economic activity. Economic intuition suggests a very simple but interesting relationship between interest rate cycles and business cycles, implying that inverted yield curves predict recessions, while upward-sloping yield curves predict economic expansions.

During the 1970s interest in asset prices as predictors of future economic activity has significantly increased. Later, in the 1990s, followed intensive research into the predictive capabilities of the yield spread for future economic activity. Stock and Watson [29], Estrella and Hardouvelis [10], Davis and Fagan [9], Bernard and Gerlach [3] and Hamilton and Kim [21], among others, find a significant positive relationship between the two and, thus, recommend the term spread as a leading indicator for future economic activity. Some of the more recent work, like, for example, Smets and Tsatsaronis [28], emphasises, however, that the relationship between the slope of the yield curve and output growth is not as well-established as had been pointed out by earlier studies. It has been suggested that in countries, which introduce tight controls on inflation, the predictive power of the term spread decreases considerably. Furthermore, the term structure loses its predictive content especially in countries in which there are irrational 'inflation scares'.

The ability of the yield curve to predict future output growth comes from a seemingly straightforward application of the expectations hypothesis of the term structure. If monetary authorities give great emphasis to interest rate smoothing, the yield spread can be represented as the aggregated change in future monetary policy rates. Interest rate smoothing has been shown to be a necessary condition for the yield spread to have predictive content for future output growth. In particular, the ability of the term spread to forecast the future path of output is probably not due to real shocks that change the real interest rate. Rather, monetary policy seems to play the central role. The term structure reflects investors' expectations about monetary policy reactions to the general state of the economy and, in particular, to supply and demand shocks. During recessions term premia of long-bonds tend to be high and policy rates tend to be low The resulting upward-sloping yield curve predicts higher growth in future. During booms, instead, interest rate hikes by the monetary authorities, and lower term premia on long-bonds, will flatten the yield curve, or might, eventually, invert it, indicating contractions in future GDP.

The predictive power of the yield curve for future output growth has been widely studied. Yet, there has been few work analysing the predictive content of the yield spread in Central and Eastern European (CEE) accession countries. Research on the dynamics of output, inflation and interest rates in these countries, which, after having abolished the central planning of their economies, first had to go through a transitional phase of market liberlisation, is still in its early stages and in the phase of being developed. This makes the accession countries obvious candidates for further research on the predictive content of the yield spread. This kind of study has important implications for policy makers in transition economies as well as for investors in the capital markets of these countries.

Our contribution is to analyse the predictive content of the yield spread and its components for future output growth in the accession countries Bulgaria, Czech Republic, Hungary, Poland, Slovakia, and Romania. The methodology we apply is an Expectations Theory-based decomposition of the yield spread into an expectations-related and a term premium component in a VAR framework. We examine the predictive content of the two components of the yield spread for future real economic activity using a model of the government bond yield spread as well as employing a model of the interbank yield spread.

Due to the implications that the characteristics of the monetary policy regime have for the predictive content of the yield spread, we put our results into the context of the conduct of monetary policy in the accession countries. Tho this end, we analyse the monetary policy reaction functions in order to check how policy rates in the CEE countries can be predicted using a Taylor rule.

2 Literature

Fama [13] investigates the usefulness of the yield curve for forecasting economic activity in a seminal paper. Inspired by the work of Fama and others, a broad strand of empirical evidence, suggesting that the yield curve is capable of predicting output growth, inflation and the timing of recessions in some countries, emerged.

In particular, Peel and Taylor [27] use the yield curve as a leading indicator of future economic activity and Estrella and Mishkin [11] point out that for forecast horizons of more than two quarters the yield spread outperforms the

US Commerce Department's index of leading indicators. This is confirmed by Estrella and Hardouvelis [10], Feroli [18] and others. Fama [15] and Mishkin [25] show that term spreads between long-term and short-term interest rates contain information about future inflation rates in the US. Haubrich and Dombrosky [20], too, find evidence that the yield curve has predictive power for real GDP growth in the US and Smets and Tsatsaronis [28] confirm this result for the US and for Germany.

However, Davis and Fagan [9] suggest more caution regarding the ability of the yield curve to forecast inflation in EU countries and Bernanke [2] finds mixed results for the predictive power of the yield curve for several economic variables in the US. This power seems to have faded away through time up to 1990 due to changes in Federal Reserve operating procedures and financial market innovations.

Concerning the prediction of recessions, in the US the yield curve was capable of predicting all five recessions that have occurred since 1960. Bernard and Gerlach [3] conclude that the yield curve can predict recessions in the US, Germany and Canada. Estrella and Mishkin [11], using financial variables for the US, Peel and Taylor [27] for the US and the UK, and Smets and Tsatsaronis [28] for the US and Germany, find similar results.

Feroli [18] emphasises the role of monetary policy in interpreting the yield spread. Like some other authors he underlines that the yield spread contains information beyond the pure policy stance of the central bank and that the functional form of the monetary policy reaction function is crucial for the information content of the yield spread for future economic activity.

Inspired by the apparent importance of expected future policy rates for the predictive power of the yield spread, Hamilton and Kim [21], Ang et. al. [1] and Favero et. al. [17] introduce and develop the decomposition of the yield spread into an expectations-related and a term premium component. Doing a real-time decomposition for the US, Favero et. al. [17] find that only the term premium component of the yield spread has a significant effect on future output growth. Hamilton and Kim [21] find that both components are significant while Ang et. al. [1] find that only the expectations-related component is significant. The explanation for the three contrasting results might be found in the different methodological approaches employed. Favero et. al. [17] decompose the yield spread in real-time and, therefore, include only historically available information in the agents' information set regarding expectations on future short-term rates, whereas Hamilton and Kim [21] and Ang et. al. [1] estimate their models on the full sample and then make in-sample forecasts, therefore, giving agents more information than they could actually have on hand in real-time.

3 Why Does the Yield Curve Predict Economic Activity?

A broad strand of empirical literature demonstrates that the yield curve can predict future inflationary dynamics, consumption, investment, recessions and probably the rate of GDP growth for different combinations of short-term and long-term interest rates and countries. The yield curve simply represents the term structure of interest rates formalised by the expectations hypothesis of the term structure, plus a time-varying term premium, and, thus, embodies market expectations of future nominal interest rates.

One needs to distinguish between the fact that, on the one hand, the yield curve directly causes changes in future output growth and, on the other hand, it predicts changes in future output growth. The direct power of the yield spread over future economic activity stems from the fact that, say, a hike in the policy rate raises short-term rates more than it does raise long-term rates and, while the yield curve flattens, investment, consumption and real growth slow down through the aggregate demand channel of monetary transmission.

The term spread has, however, been found to have considerable power in forecasting output growth because it partly reflects investors' expectations about monetary policy reactions to supply and demand shocks apart from the causal link between a steeper yield curve and higher economic growth. Anticipating the counter-cyclical monetary policy of the monetary authorities, which in turn is related to future economic activity, investors bring the yield curve in line with the expected path of future policy rates. Smets and Tsatsaronis [28] which analyse the economic determinants of the term structure to answer the question why the yield curve predicts economic activity in general and the path of output growth in particular, suggest that the predictive content of the yield curve depends on the variables in the information set of the monetary authorities and how they react to them. The conclusion that the predictive content of the yield spread critically depends on the the specification of the monetary authorities' monetary policy reaction function is drawn by Feroli [18], too.

The predictive power of the yield spread is contingent on the functional form of the monetary policy reaction function and, in particular, on interest rate smoothing as Feroli [18] and others point out. In the presence of interest rate smoothing the one-period change in the policy rate becomes the instrument of monetary policy and since the yield spread can be expressed as the accumulated differences in the policy rate during the life-time of the long-term bond used in the construction of the spread, the central bank's expected future path of output is, in some sense, contained in the yield spread. In very simplified form, the yield spread $S_t^{2,1}$ between a two-period and a one-period bond in equations 1 and 2 can be expressed as the difference between the expected policy rate $E_t i_{t+1}^1$ in period $t+1$ and the actual policy rate i_t^1 in period t. By this token, the yield curve is shaped as it is because it responds to the expected future dynamics of key economic variables. The yield spread extracts expected inflation and output from the expected path of interest rates, which is determined by the central bank's monetary policy reaction function. Therefore, the specification of the monetary policy reaction function will move centre stage in the next subsection.

$$S_t^{2,1} = i_t^2 - i_t^1 \tag{1}$$

$$S_t^{2,1} = \frac{1}{2}(E_t i_{t+1}^1 - i_t^1) \tag{2}$$

An important contribution to the predictive content of the yield spread in differing monetary regimes comes from Bordo and Haubrich [4] who find that the predictive content in the US varies with different degrees of credibility of the monetary authorities and with different degrees of inflation persistence. High credibility of the monetary authorities can be associated with a low predictive content of the yield curve because in high-credibility monetary environments, such as a gold standard, inflationary or real shocks push up only short rates leaving long rates unchanged. Obviously, in this case a flatter or inverted yield curve does not convey information about lower future growth, as it would have done instead in a laxer monetary policy environment, but rather sends out a noisy signal about future GDP growth.

Stressing, instead, the importance of the term premium component of the yield spread Favero et. al. [17] deliver a different explanation of the leading indicator property of the yield spread. They find a structural interpretation for the interaction between the output gap and the term premium in the form of an aggregate-demand type intertemporal Euler equation incorporating the degree of agents' risk aversion which is supposed to be time-varying. In times of particularly high risk-aversion and term premia, monetary policy seems to have less power in determining output fluctuations than during times in which risk aversion is low.

Concerning the ability of the yield curve to predict future inflation, we assume for a second that the yield curve were made up of only two securities with two different maturities - short and long. The nominal short bond rate i^1 corresponds to the real interest rate r plus expected inflation $E\pi_1$ during the life-time of the short bond:

$$i^1 = r + E\pi_1 \tag{3}$$

If neither the real interest rate r nor the inflation rate π_2 are expected to change over time, the nominal long bond rate i^2 will be identical to the nominal short-bond rate i^1. By the expectations hypothesis, the nominal long-term yield i^2 is one half the short-bond yield in period 1, i^1_1, plus one half the expected short-bond yield in period 2, i^1_2:

$$i^2 = \frac{1}{2}(i^1_1 + i^1_2) \tag{4}$$

In the absence of a term premium, the nominal rate of return of a security of any specific maturity is equal to the real interest rate plus the expected inflation rate for the maturity under consideration. For our long bond rate i^2 this implies that it will be equal to the real rate r plus the expected inflation rates $E\pi_1$ and $E\pi_2$:

$$i^2 = r + \frac{1}{2}(E\pi_1 + E\pi_2) \tag{5}$$

This implies for the slope of the yield curve:

$$i^2 - i^1 = \frac{1}{2}(E\pi_2 - E\pi_1) \tag{6}$$

Clearly, the slope of the yield curve gives us information about the expected change in the inflation premia inherent to the long bond. If long-term inflationary expectations $E\pi_2$ rise, the yield of the security with the long maturity has to rise and if at the same time the real interest rate and inflationary expectations regarding the short maturity do not change, the slope of the yield curve will increase. Hence, the steepening of the yield curve conveys information on expected future inflation. The introduction of a variable term premium could complicate the analysis but does not nullify the informational content of the yield curve.

Obviously, the information content of the yield curve for future economic activity depends as well on the type of shock that hits the economy and leads to changes in the yield curve. If a small open economy is hit by an anticipated demand shock, e.g. a change in tax laws inflation will gradually rise and after a restrictive central bank reaction inflation and output converge back to equilibrium. Market expectations of the central bank's reaction and the reaction itself drive the development of the term spread. The expected demand shock first raises the term spread and then after the shock the term spread tightens due to hikes in the policy rate. Within the context of the anticipated demand shock, the term spread has positive predictive power for future real economic activity as well as for future inflation.

Anticipated supply shocks, too, first lead to expectations about pre-emptive monetary policy, accompanied by an expected rise in inflation and an inflation-driven strong appreciation of the real exchange rate. This leads to a fall in real output even if the movement of the term spread is similar to the one in the demand shock case because agents' perceived reaction function puts strong emphasis on inflation stabilisation. So the term spread's ability as an indicator of future inflation remains unchanged in the advent of a supply shock, but the term spread is no longer able to indicate future real economic activity.

Regarding agents' expectations on recessions, the response of the yield curve will depend on the expected size and duration of the recession's effect on short-term interest rates. Clearly, short-term interest rates are expected to decline during recessions. This could be due as well to an expected reduction of real rates of return during a recession which has nothing to do with the expectations hypothesis. However, during recessions markets expect lower inflation rates and according to the expectations hypothesis longer-term interest rates should decline on the grounds of lower inflationary expectations and lower expected short term rates so that the yield curve is likely to flatten or will even become inverted prior to recessions. Once again, term premia could decrease the information content inherent to the yield curve.

One of the most important findings of the more recent literature on the predictive content of the yield spread is that the yield spread contains information for future output growth beyond current and lagged output and the information on the monetary policy stance as, e.g., Estrella and Hardouvelis [10] and Hamilton and Kim [21] point out. Nearly all authors find that expectations of future monetary policy which are not captured in the policy rate or in other indicators of current monetary policy play a crucial role in the spread's predictive abilities. These expectations might include agents' views on term premia, inlation premia or other factors influencing long bond yields. If investors expect higher future money supply growth they revise expectations about the future real rate downwards and, as a consequence, future output will expand. If the increase in the inflation premium due to higher expected money supply growth will now exceed the decrease in the future real rate the yield curve will steepen, thus, creating a positive correlation between the slope of the yield curve and future changes in real output.

Therefore, in the next subsection we will elaborate on the conduct of monetary policy, the formation of expectations about future monetary policy and how they influence the predictive power of the yield spread.

3.1 The Specification of Monetary Policy Reaction Functions and the Predictive Power of the Yield Spread

The ability of the yield spread to predict output fluctuations has been shown to be contingent on the monetary authority's reaction function. In particular, expectations on monetary policy actions are crucial for the spread to predict future output. For the US, Feroli [18] suggests that the post-1979 decrease in the yield spread's predictive power is due to a shift in the monetary policy reaction function at that time.

Usually it is portrayed that a rise in the short-term rate will flatten the yield curve and slow down real growth. Thus, a flatter yield curve precedes a slowdown in output. When, however, the short rate or other measures of the monetary policy stance are included into an output forecasting equation, the yield spread will, nevertheless, retain its significance. This means that the yield spread contains information beyond the pure monetary policy stance already contained in the short rate.

Feroli [18] shows that the predictive power of the yield spread is contingent on the functional form and the parameter values of the monetary authority's policy reaction function. Two elements are crucial for the spread to contain information about future output. Firstly, interest rate smoothing and, secondly, counter-cyclical monetary policy. In the presence of interest rate smoothing, the difference in one-period rates becomes an instrument of monetary policy. The difference in one-period rates represents the central bank's perception regarding the future path of output and inflation. These expectations are contained in the yield spread, which, with the help of the expectations hypothesis, can be expressed as the difference of the one-period rates. If agents are aware of central bank interest rate smoothing they incorporate their expectations on future inflation and output growth into the yield spread.

The structural parameters related to the conduct of monetary policy vary over time and are the second force explaining the predictive content of the yield spread. Clarida et al. [8] and Estrella et al. [12] show that these structural parameters changed significantly after Paul Volcker became Chairman of the Federal Reserve in the third quarter of 1979. The anti-inflationary monetary policy stance during the chairmanship of Volcker could have contributed to the declining predictive content of the yield spread.

In theory, incorporation of the expectations model of the term structure makes long-term interest rates equal to average expected future short-term rates. Hence, if short-term rates are defined by a process which can be proxied by a Taylor rule, the parameters which define the relationship between the yield spread and future output dynamics are governed by the coefficients of the monetary policy reaction function. A short-coming is, however, that a model assuming the validity of the expectations hypothesis right from the start is prone to neglect the importance of risk premia.

Most macro models including the expectations model assume that the pure expectations hypothesis will hold and that the central bank policy decisions can be approximated by a Taylor rule.

$$i_t^N = \frac{1}{N} E_t(i_t^1 + i_{t+1}^1 + i_{t+N-1}^1) \tag{7}$$

$$i_t^1 = \rho i_{t-1}^1 + (1-\rho)[i^* + \beta(E_t \pi_{t,k} - \pi^*) + \gamma E_t y_{t,q}] \tag{8}$$

According to the pure expectations hypothesis, the yield of the long bond i_t^N is a function of the sequence of one-period interest rates i_t^1 to i_{t+N-1}^1. The forward-looking version of the Taylor rule incorporates the preferences of the central bank for interest rate smoothing and the stabilisation of inflation and output. i^* and π^* are the monetary authority's target nominal interest rate and inflation rate, $E_t \pi_{t,k}$ and $E_t y_{t,q}$ are, respectively, the expectations of the inflation rate k periods ahead and the output gap q periods ahead. ρ is a parameter measuring the monetary authority's preference for interest rate smoothing.

If the central bank gives great emphasis to output stabilisation the yield spread will probably be a good indicator for predicting future output, if it gives great emphasis to inflation targeting the yield spread has more informational content regarding future inflation. In small macro models including an aggregate demand equation, the Phillips curve, the rational expectations hypothesis and a monetary policy reaction function, a central bank putting great emphasis on the inflation gap will hike rates in reaction to supply shocks like , for example, tax increases. If the supply shock is anticipated by agents, long rates will most likely go up already some time before the shock in anticipation of higher future policy rates and the yield curve steepens in anticipation of the supply shock and in anticipation of a central bank which can be expected to raise rates for disinflationary purposes. In the context of a supply shock coupled with a disinflationary bias of monetary policy the steepening of the yield curve represents expectations about higher future inflation.

On the other hand, in a setting with a central bank which has the principal objective of stabilising the output gap agents expect the central bank to cut rates in order to prevent any negative impact of the supply shock and the yield curve will most likely become flatter or even invert. The example of an adverse supply shock illustrates how the predictive power of the yield spread might be dependent on the parameters of the central bank's monetary policy reaction function. In the case of a supply shock coupled with a central bank which gives great weight to output stabilisation the flatter yield curve represents lower expected future output growth.

As the yield spread does retain its significance for future output growth even if a short-term interest rate is included into a regression of future output growth on the yield spread and on inflation, expectations on future monetary policy, not being captured in the fed funds rate or in other available indicators of current monetary policy, play a crucial role in the spread's predictive abilities. Risk premia could play a central role here, as was pointed out by Favero et al. [17]. In times of heightened risk aversion the information content of monetary policy for future output growth could be more subdued, while in turn the information content of risk premia could be higher. This should be especially the case in countries in which risk premia do really matter.

Therefore, we include a measure of the term premia along the lines of Campbell and Shiller [5] in our analysis on the relationship between the yield spread and future output growth. A model trying to analyse the predictive content of the yield spread for CEE countries needs as well to pick up from the work of Feroli [18] and, thus, needs to analyse in as much the predictive content in CEE countries depends on the specification of the monetary reaction function.

Therefore, it will be of prime importance to estimate the CEE monetary policy reaction functions and compare them to classical Taylor rules.

4 Data

As the availability of interest rate quotations varies across countries we have to use various time spans for the different countries. This is especially true for government bond and treasury bill data but less so for interbank interest rate data.

The government bond and the treasury bill interest rate data is retrieved from DataStream and from the respective national central banks' websites. The data-sets for interbank rates come from DataStream. The interbank series used are the analogue of LIBOR (London InterBank Offered Rates). Apart from the five-year government bond yield only data for the shorter end of the term structure are used (the tenor with maturities from 1 month to 1 year). There are a few reasons for such a restriction. The markets in transition countries are often imperfect and non-developed and the interest-rate market is not an exception. Although the countries under consideration have the most developed markets among the transition countries, they are still not at the level of developed countries. Banking institutions lend and borrow mostly money with the shortest maturities and official quotations of interest rates often exist only for maturities up to 1 year. The pricing of instruments with longer maturities is based on swaps and rates calculated from swaps. These rates are quoted by DataStream for some countries, but as mentioned above, with longer maturities the market is even more imperfect. Mostly there exist only government bonds with higher maturities; municipal or corporate bonds exist only rarely. So the reasons for restricting the data-sets to maturities up to one year are market imperfections and liquidity (no trading with longer maturities - the volume of trades is often zero for longer maturities) and data availability. These interbank rates, although in general not risk-free (there is always a risk of bankruptcy of a bank, which is incorporated in the rates), are used by banks and other financial institutions as the lending and borrowing rates, and therefore they are used for the pricing of derivatives.

Data on GDP, CPI inflation, exchange rates and eurozone and German interest rates is retrieved from DataStream, too.

5 Decomposition of the Yield Spread

The usefulness of the decomposition of the yield spread into a term reflecting future monetary policy and a term reflecting the term premium, in order to understand why yield curve fluctuations help to predict subsequent economic activity has been underlined by work such as Hamilton and Kim [21], Ang et al. [1] and Favero et al. [17].

When the expectations theory of the term structure is augmented by a time-varying term premium, the long rate i_t^n equals the sum of the weighted expected future short rate changes plus the term premium:

$$i_t^n = \frac{1}{n}\sum_{j=0}^{n-1} E_t i_{t+j}^m + TP_t \tag{9}$$

Then, the yield spread between the long rate i_t^n and the short rate i_t^m can be written as

$$i_t^n - i_t^m = \left(\frac{1}{n}\sum_{j=0}^{n-1} E_t i_{t+j}^m - i_t^m\right) + TP_t \tag{10}$$

substituting for the term premium the yield spread can be rewritten as

$$i_t^n - i_t^m = \left(\frac{1}{n}\sum_{j=0}^{n-1} E_t i_{t+j}^m - i_t^m\right) + \left(i_t^n - \frac{1}{n}\sum_{j=0}^{n-1} E_t i_{t+j}^m\right) \tag{11}$$

Carrying out the decomposition of the yield spread, important assumptions on how agents form expectations on future interest rates have to be made. Different assumptions on the way agents form interest rate expectations can cause quite large differences in the results of the decomposition and in the resulting expectations-related component and the term premium of the yield spread. This can be illustrated, for example, by comparing the work of Hamilton and Kim [21], Ang et al. [1] and Favero et al. [17]. While Favero et al. [17] and Ang et al. [1] find that only the term premium plays a role for future output growth, Hamilton and Kim [21] find that both the expectations-related and the term premium component contain some predictive content for future GDP growth. Their very different results are partly due to different assumptions about the formation of expectations by the agents. While Favero et al. [17] chose a real-time approach in which agents have only historical information available, Hamilton and Kim [21] and Ang et al. [1] use ex-post data for the decomposition of the yield spread, giving agents more information than they realistically would have when forming expectations.

5.1 Hamilton and Kim [21] Decomposition

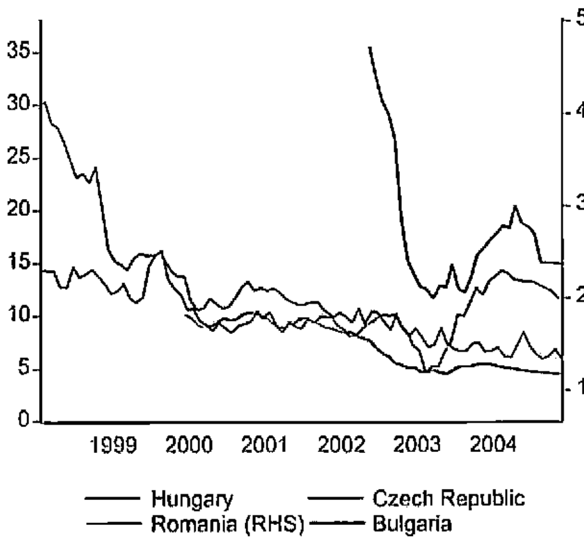
Hamilton and Kim [21] find that both the expectations-related (ER) and the term premium (TP) components are significant in forecasting future output growth. While they find that interest rate volatility is correlated with the term structure and GDP, interest rate volatility does not seem to account for the yield spread's usefulness for predicting GDP growth. Both the ER and TP components are important for predicting future output at short horizons. At long horizons, however, the ER component plays a much more dominant role. They show that the yield spread has information content for future GDP growth beyond the information contained in lagged GDP growth and in the monetary policy stance.

Hamilton and Kim [21] use ex-post observed short-rates instead of ex-ante expected rates. They use perfect foresight and rational expectations and, thus, assume that the forecasting errors of the agents will even out over time. We adapt Hamilton Kim to monthly data and define the yield spread by the difference between the yield of five-year government bonds and three-month treasury bills. We fit the following equation which relates the twelve-months ahead real growth of log GDP to the ER and TP components of the yield spread

$$\Delta_k y_{t+k} = \alpha_0 + \alpha_1 \left(\frac{1}{n} \sum_{j=0}^{n-1} E_t i_{t+j}^m - i_t^m \right) + \alpha_2 \left(i_t^n - \frac{1}{n} \sum_{j=0}^{n-1} E_t i_{t+j}^m \right) + u_t \quad (12)$$

We estimate the model as Hamilton and Kim [21] do with the instrumental variables of a constant, the three-month treasury bill yield and the five-year government bond yield. Therefore, apart from perfect foresight about future short-rates we assume a rational expectations error term u_t which will be uncorrelated with any variable known at time t. With the Hamilton Kim decomposition we lose the last five years of historical data for the estimation of real GDP growth because market expectations on the interest rate and on the term premium are calculated on ex-post data. Therefore, in order to be able to calculate the term premia up to 2004 and not only up to the year 1999, we forecast three-month interest rates from a VAR according to Campbell and Shiller [5] until the year

Figure 1: Term premia between five-year government bonds and three-month treasury bills according to Hamilton-Kim Decomposition



2009 in order to calculate the Hamilton Kim term premia up to the end of 2004.

Estimation results from the Hamilton Kim decomposition in equation 12 are displayed in table 1. We drop Romania from the analysis because we have too few observations at hand.

According to the Hamilton Kim decomposition, the expectations-related and the term premium component of the yield spread predict future real GDP growth only in the Czech Republic. In Hungary and Bulgaria, instead, the coefficients come out hardly significant and the magnitude of the coefficients is so small that the impact of the expectations-related and the term premium component on future GDP growth can be neglected. This is in contrast to the Czech Republic where a one percent rise in the expectations-related component predicts an increment of one-year-ahead GDP of roughly 70 basis points, while a one percent rise in the term premium component forecasts a GDP increase of approximately half a percentage point.

These are, by all standards, significant GDP changes. However, the Hamilton Kim decomposition suffers from the fact that the yield decomposition is carried out on ex-post interest rate data, thus giving agents more information

Table 1: Regression results for Hamilton Kim decomposition of the yield spread

	Bulgaria	Czech Republic	Hungary
α_0	-0.122 (-0.1)	4.885 (12.9)	0.456 (9.64)
α_1	-0.066 (-5.1)	0.685 (5.56)	-0.027 (-1.6)
α_2	-0.003 (-1.0)	0.524 (3.77)	-0.035 (-1.9)
R^2	0.28	0.62	0.10
$D.W.$	0.88	0.23	0.34

Sample: Jan'97 - Sep'04, t-statistics are reported in parentheses

than they realistically could have on hand when making their investment decisions.

5.2 Ang et al. [1] Decomposition

Ang et al. [1] advocate a model using a very limited number of factors to summarise the information contained across all maturities of the yield curve. The three factors in their Gaussian Vector Autoregression are the short rate as proxy for the level of the yield curve, the yield spread as proxy for the slope of the yield curve and the rate of real GDP growth as a macroeconomic factor. They project expectations for the short-term rate, the yield spread and GDP estimating their VAR on the full sample and calculating the expected realisations of the variables by simulating the VAR forward, thus, giving agents information which they would not have had available in real time. Only the nominal short rate and inflation are found to be important for forecasting future GDP growth, while the term premia seems to matter less. The maximal maturity difference is identified as the best measure of slope in their context. They suggest, however, that the nominal short rate contains more information about GDP growth than any yield spread, which is in stark contrast to the evidence from unconstrained OLS regressions. This could be due to their chosen forecast period ranging from 1990 to 2001. Other work, e.g. Bernanke [2], has shown too that the predictive content of the term spread is not constant over time and that it probably declined during the 1990s. Favero et al. [17], covering a much longer time span, find a much more important role for the term premia. This is, however, not only due to the different sample period but as well to big differences in the methodology chosen.

5.3 Favero et al. [17] Real-time Decomposition

Favero et al. [17] underline the usefulness of the decomposition of the yield spread, in a term reflecting future monetary policy and a term reflecting the term premium, for understanding how yield curve fluctuations help to predict future economic activity. They decompose the yield spread in an expectations-related component and in a term premium component updating the information set of agents in real-time. They identify the best forecasting model for output as a model containing the term premium, the short-term interest rate and inflation and find a structural interpretation for their results which derives from investor's high risk aversion during times of heightened uncertainty.

If inflation surprises do not cancel out over time, ex-post observed returns can not be used as a proxy for ex-ante returns. Therefore, in contrast to Hamilton and Kim [21] and Ang et al.[1], Favero et al. [17] do a model in real time in which investors can use only historically available information to generate their predictions.[1] For each point in time they estimate a VAR in the three-month treasury bill yield, the spread of the five-year government bond yield over the three-month treasury bill yield, the inflation rate and the yearly change in real GDP and then simulate the rolling VAR forward, using only historical data to calculate the expectations-related and the term premium component. They find that the nominal short-rate, inflation and the term premium component of the yield spread are the most important variables predicting future GDP growth.

5.4 VAR Decomposition

What we will call the VAR decomposition is the decomposition of the yield spread into an expectations-related component, calculated on the basis of a VAR forecast of future short-term rate changes according to Campbell and Shiller [5], and a residually determined term premium component.

The specification of our VARs includes the yield spread between five-year government bond yields and three-month treasury bill yields, the monthly change in three-month treasury bill yields and the yield spreads of the respective country's treasury bills over the yield of German treasury bills to control for the influence of foreign interest rates on local short-term rate fluctuations.

After estimating the VAR on the full sample ranging from 1995 to 2004 we calculate the expectations-related component of the term spread from the

[1] With the exception of GDP data. For real GDP they use revised data.

expected future short-term rate changes of in-sample forecasts using equation 14. This means we have to concede agents more information than they actually could have on hand when forming expectations on future interest rates.

If the expectations theory gave an accurate picture of the true term structure of interest rates, the actual spread $S_t^{(n,m)}$ would be equal to the expected theory-based spread $E_t S_t^{(n,m)^*}$:

$$S_t^{(n,m)} = E_t S_t^{(n,m)^*}, \tag{13}$$

$$\text{where } S_t^{(n,m)^*} = \sum_{i=1}^{k-1} \left(\frac{k-i}{k} \right) \Delta^m r_{t+im}^m \tag{14}$$

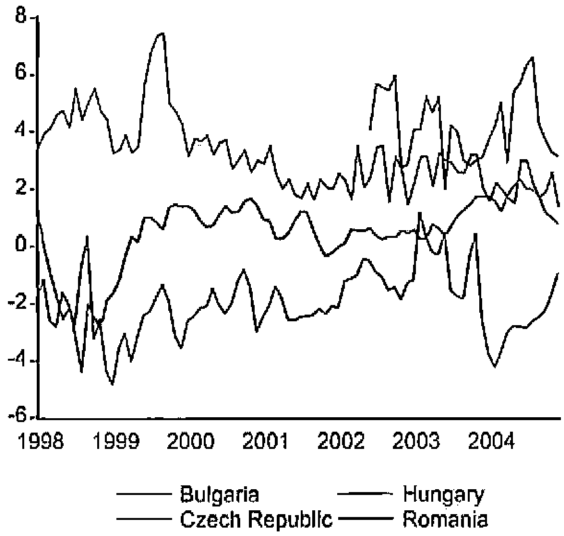
$S_t^{(n,m)^*}$ is the theoretical spread between the long-term and the short-term bond, where Δ^m stands for a change in the short-rate measured over m periods. The theoretical spread would prevail if agents had perfect foresight about future interest rates.

We do as well a VAR with a rolling estimation window so as to give agents only historical information (apart from revised data for GDP, however) for forecasting monthly short-rate changes. In this case we make use of a 48 month rolling window and estimate the VAR from the first window November 1995 - December 1999 up to the last window January 2000 - December 2004. After each estimated window we calculate the theoretical spread and the term premium according to equation 14.

The VAR is estimated on the sample from November 1995 to December 1999 and the forecast of short-rate changes is then carried out for the period between January 2000 and March 2004. Then we calculate the theoretical spread and the term premium based on the forecast data.

Afterwards we employ the seven forecasting models Favero et. al. [17] have suggested for the US and fit them for the CEE countries. Favero et al. [17] identify the dominating model for forecasting GDP growth as a model including the yield spread, decomposed into an expectations-related and a term premium component, the short-term nominal rate and inflation. Their model resembles an aggregate demand model with a short-term real interest rate. It will be interesting to see whether this benchmark model delivers a good description of the predictive content of the term spread in the accession countries.

Figure 2: Yield spread of five-year government bonds over three-month treasury bills

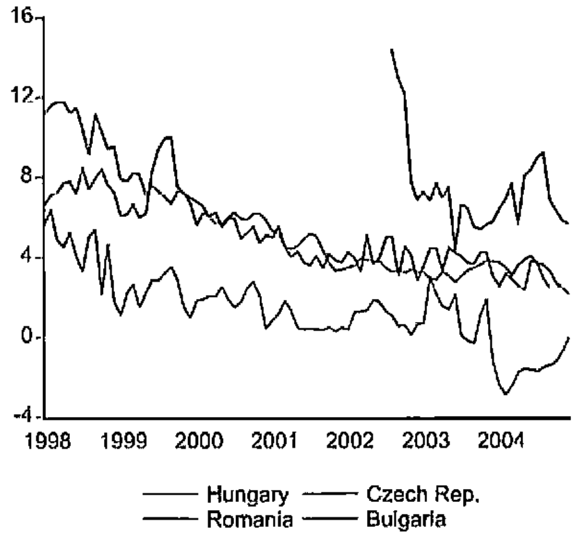


5.4.1 Term Premia of Five-year Government Bonds

Where available, we use the yield spread of five-year government bonds over three-month treasury bills in order to investigate the information content in the yield spread. This spread has been shown to be significant in various studies on various countries. For the Czech Republic and Hungary we measure the risk premium of the yield spread of five-year government bonds over three-month treasury bills. In the case of Bulgaria and Romania the central banks provide only interest rates for treasury bills and treasury bonds of an average maturity, which we use to calculate the yield spread. No adequate data to carry out the analysis was available for other major accession countries such as Poland and Slovakia.

The results for the term premia and the expectations-related component are presented in figures 3 and 4 for the sample from January 1998 to December 2004. While our estimation sample is 1995 to 2004 we display only the period between between 1998 and 2004 in the two figures. This is for higher visibility because the TP and ER components of some countries were extremely high in the period 1995-1997.

Figure 3: Term premia of government bonds ($5y - 3m$) yield spread

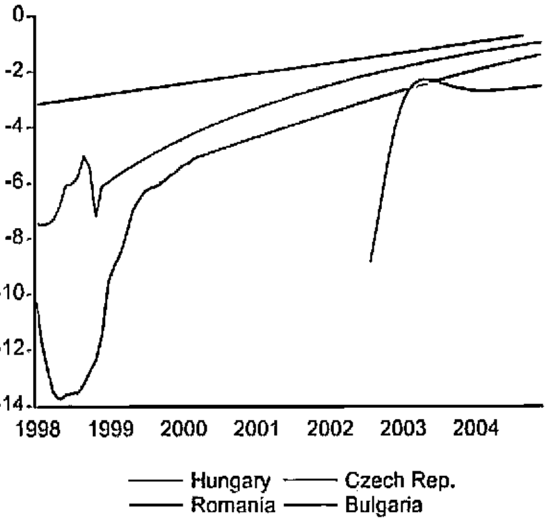


Not surprisingly, the term premium for Romania, for which, however, we have only a very short time series, is the highest. The term premia for the other three countries are in a more normal range. The Hungarian term premia heads into negative territory during the year 2003 in which Hungary experienced a speculative attack, speculation about the postponement of its EMU entry-date and much higher interest rate volatility than the other countries (see appendix B for details about the Hungarian situation during 2003).

The expectations-related components of the yield spread in figure 4 indicate that in all four countries the theory-based yield spread was actually negative, and yield spreads in Romania, Bulgaria and the Czech Republic were only positive due to the large term premium component. The dynamics of the negative expectations-related component between 1998 and 2004 probably stands for a mitigation of the expectations of lower interest rates. At the beginning of the period interest rates were still very high and agents expected rates to fall markedly during the convergence process of the accession countries. With the continuing normalisation of inflation and monetary policy in these countries the expectations-related term approached zero.

Favero et al. [17] identify the dominating model for forecasting GDP growth

Figure 4: Expectations-related component of $(5y - 3m)$ yield spread



as a model including the yield spread, decomposed into an expectations-related and a term premium component, the short-term nominal rate and inflation. We replicate all of their seven models in order to identify the most adequate model for the CEE countries. Equation 1 contains the yield spread only, equation 2 contains the spread and the short-rate, equation 3 contains the spread, the short rate and inflation, equation 4 contains the short rate and inflation and equation 5 contains the long rate and inflation. Equations 6 and 7 contain the expectations-related component ER_t^{CS} and the term premium component TP_t^{CS} of the yield spread, which, however, in contrast to Favero et al. [17] are not calculated in real-time but according to Campbell and Shiller [5] with a VAR estimated on the full sample as described above. In equation 6 the expectations-related and term-premium components are augmented by the short-rate and by inflation. Equation 8 instead estimates output growth using the decomposition of the yield spread with ex-post data suggested by Hamilton and Kim [21]. All models contain a constant. The specification of equations 1 to 8 is as follows:

where $\Delta y_{t,t+12}$ is the change in log real GDP in the upcoming twelve months, i_t^{3m} is the yield on three-month treasury bills, i_t^{5y} is the yield on five-year government bonds and $S_t^{(5y-3m)}$ the yield spread between five-year government bonds

Table 2: Forecasting models for future GDP growth

$Eq.1$	$\Delta y_{t,t+12} = \alpha_0 + \alpha_1 S_t^{(5y-3m)} + u_t$
$Eq.2$	$\Delta y_{t,t+12} = \alpha_0 + \alpha_1 S_t^{(5y-3m)} + \alpha_2 i_t^{3m} + u_t$
$Eq.3$	$\Delta y_{t,t+12} = \alpha_0 + \alpha_1 S_t^{(5y-3m)} + \alpha_2 i_t^{3m} + \alpha_3 \pi_t + u_t$
$Eq.4$	$\Delta y_{t,t+12} = \alpha_0 + \alpha_1 i_t^{3m} + \alpha_2 \pi_t + u_t$
$Eq.5$	$\Delta y_{t,t+12} = \alpha_0 + \alpha_1 i_t^{5y} + \alpha_2 \pi_t + u_t$
$Eq.6$	$\Delta y_{t,t+12} = \alpha_0 + \alpha_1 i_t^{3m} + \alpha_2 \pi_t + \alpha_3 ER_t^{CS} + \alpha_4 TP_t^{CS} + u_t$
$Eq.7$	$\Delta y_{t,t+12} = \alpha_0 + \alpha_1 ER_t^{CS} + \alpha_2 TP_t^{CS} + u_t$
$Eq.8$	$\Delta y_{t,t+12} = \alpha_0 + \alpha_1 ER_t^{HK} + \alpha_2 TP_t^{HK} + u_t$

Table 3: Correlations between ER, TP and (5y-3m) yield spread

ρ	*Bulgaria*	*CzechRep.*	*Hungary*	*Romania*
ER_t^{CS}, TP_t^{CS}	−0.17	−0.95	−0.83	−0.86
$ER_t^{CS}, S_t^{(5y-3m)}$	−0.17	0.40	0.38	−0.30
$TP_t^{CS}, S_t^{(5y-3m)}$	0.99	−0.09	0.20	0.75

and three-month treasury bills, π_t is the inflation rate and ER_t^{CS} and TP_t^{CS} are the expectations-related and the term premium components of the yield spread. The results of the regressions are displayed in table 4.

As shown above, the term premia is residually constructed from forecast short-rate changes. Therefore, by construction there should be a highly negative correlation between the expectations-related and the term premium components of the yield spread. In fact this is true for the Czech Republic, for Hungary and for Romania as can be seen from table 3. However, the decomposition does not seem to work for the case of Bulgaria, probably due to the extreme jumps in the Bulgarian term spread in the years 1996 and 1997.[2]

The yield spread seems to work quite well as a predictor for one-year ahead output growth in the Czech Republic as can be seen from equations 1, 2 and 3. For Bulgaria and Hungary instead the yield spread is not generally significant and in the cases in which it comes close to conventional levels of significance it seems to predict contractions in future GDP rather than economic expansions. For Romania our time series is too short and the resulting significance levels are

[2]The Bulgarian yield spread was 4176.4 percent in October 1996. For the period from January 1998 to December 2004 the correlation between the expectations-related component and the term-premium component is -0.85 and compares well to the results for the other countries.

Table 4: GDP forecasting equations with (5y-3m) government bond yield spread

		Bulgaria		*Czech Rep.*		*Hungary*		*Romania*	
eq1	c	2.681	2.0	2.758	15.2	0.379	9.8	1.282	3.7
	S_t	0.004	1.3	1.084	8.7	-0.006	-0.4	-0.108	-1.4
	$\bar{R}^2$	0.01		0.42		-0.01		0.04	
eq2	c	-0.236	-0.2	4.791	12.8	0.509	8.1	1.349	2.3
	S_t	-0.006	-1.8	0.596	4.4	-0.036	-1.9	-0.106	-1.4
	i_t^{3m}	0.131	5.4	-0.268	-6.0	-0.015	-2.6	-0.004	-0.1
	$\bar{R}^2$	0.24		0.56		0.05		-0.00	
eq3	c	0.265	0.3	4.852	12.6	0.506	8.0	1.602	2.3
	S_t	-0.001	-0.4	0.567	4.0	-0.037	-2.0	-0.132	-1.5
	i_t^{3m}	0.041	1.7	-0.285	-5.6	-0.014	-2.2	-0.004	-0.1
	π_t	0.310	6.9	0.285	0.7	-0.026	-0.8	-0.135	-0.7
	$\bar{R}^2$	0.50		0.56		0.05		-0.02	
eq4	c	0.322	0.3	5.877	6.8	0.592	12.9	0.802	8.8
	i_t^{3m}	0.035	1.9	-0.412	-9.2	-0.012	-3.6	-0.008	-3.9
	π_t	0.315	7.3	0.762	1.8	-0.063	-2.4	-0.034	-2.0
	$\bar{R}^2$	0.50		0.50		0.28		0.37	
eq5	c	0.913	0.9	6.104	14.0	0.524	8.4	1.314	2.0
	i_t^{5y}	0.002	0.9	-0.421	-7.0	-0.011	-1.9	-0.021	-0.8
	π_t	0.351	0.4	0.771	1.5	-0.015	-0.5	-0.029	-0.2
	$\bar{R}^2$	0.49		0.33		0.03		-0.05	
eq6	c	6.685	2.0	4.827	12.3	0.551	8.2	-0.462	-0.4
	i_t^{3m}	0.066	2.5	-0.301	-5.1	-0.027	-2.8	0.120	1.8
	π_t	0.303	6.9	0.291	0.7	-0.026	-0.8	0.077	0.4
	TP_t^{CS}	-0.002	-0.7	0.568	4.0	-0.049	-2.5	-0.048	-0.5
	ER_t^{CS}	3.301	2.1	0.543	3.7	-0.076	-2.7	0.201	1.1
	$\bar{R}^2$	0.52		0.55		0.07		0.09	
eq7	c	-0.915	-0.2	4.077	11.3	0.402	8.7	1.356	3.9
	TP_t^{CS}	0.004	1.1	0.754	5.4	-0.010	-0.6	-0.086	-1.1
	ER_t^{CS}	-1.703	-0.9	0.953	7.8	-0.002	-0.1	-0.033	-0.3
	$\bar{R}^2$	0.01		0.50		-0.01		0.04	
eq8	c	-0.122	-0.1	4.885	12.6	0.456	9.6	1.315	3.1
	TP_t^{HK}	-0.003	-1.0	0.524	3.7	-0.035	-1.9	-0.105	-1.3
	ER_t^{HK}	-0.066	-5.1	0.685	5.4	-0.027	-1.6	-0.103	-1.2
	$\bar{R}^2$	0.26		0.59		0.08		-0.00	
obs		94		93		92		28	

R-squared reported is adjusted R-squared too low to derive any conclusions.

The decomposition of the yield spread into expectations-related and term premium component in equations 6, 7 and 8 does not give any clear indications on the predictive content of the ER and TP components for our set of countries as a whole. For the Czech Republic equation 6 confirms the importance of the term premium but the expectations-related component is shown to be important for predicting future GDP growth, too. This, however, can not be said for Bulgaria and Hungary. In Bulgaria only the TP component is significant and predicts GDP expansions, while in Hungary both components are significant they indicate instead contractions.

A quick look on table 4 shows that the fit of the different equations is very heterogeneous for the four countries. There is no single equation which could be identified as the best model for accession countries in general. For the Czech Republic equations 2, 3, 6 and 8 have the highest fit according to the adjusted R-squared. The Hamilton Kim decomposition has the highest it suffers, however, from major methodological problems as investors' information sets contain ex-post data. According to Favero et al. [17] equation 6 has the highest forecasting predictability for the US.

With the exception of the Czech Republic we find little evidence, for the yield spread and its components to predict future economic activity. A yield spread, an expectations-related component or a term premium component of 1% all predict future GDP increases of around 60 basis points. In Hungary, if anything, the yield spread and its components seem to predict negative GDP growth. The magnitude of the predicted GDP change is however only between four and eight basis points and, hence, seems negligible. In Bulgaria, instead, we find an important role played by the expectations-related component which, when at 1%, suggests a massive increase of 3.3 percent in real GDP and a significant coefficient on inflation which predicts a 30 basis points increase in output.

Unfortunately, due to limited data availability we have to use various time spans across countries and for some countries we have only very few observations to estimate the models displayed in table 4. Therefore, it might be worthwhile to pool our data. After pooling the data we cannot make statements about the predictive content of the yield spread in any particular country, however, we have information on the predictive content in a 'typical' CEE country. We pool the data in spite of the fact that the coefficients of the countries in equation 6 of table 4 have very different slopes. In fact the Wald test in table 5 decisively rejects the identity restriction for each coefficient and the panel restriction, too. Even if the Wald test rejects the panel, the panel, however, allows us to analyse the predictive content of the yield spread in a hypothetical 'average' accession country.

The pooled regression in table 5 suggests that the ER and TP components of the yield spread, the three-month interest rate and the inflation rate all have some predictive content for future output growth in the 'typical' CEE country. A higher nominal short-term interest rate unsurprisingly predicts lower twelve-month ahead GDP growth. Observing a higher short-term rate agents expect so see slower future output growth, apart from the causal link between higher short-term rates today and slower output growth in the future. Inflation finally comes in with the correct sign in the pooled regression. Higher inflation today should indicate stronger future output growth because often it is the inflation premium which bids up the long-bond yield at the beginning of an economic expansion. In this way expectations on higher inflation during economic booms are priced in long-bond yields already before the booms actually start. The term premium component has the expected positive sign but its coefficient is much smaller than the expectations-related component's one and its significance is lower, too. The only coefficient which has a relatively high magnitude is the expectations-related component. An expectations-related component of 1% implies a 25 basis points rise in GDP within twelve months. This means that agents expect that within one year the central bank will get slightly closer to a restrictive policy move due to an improved growth outlook.

Considering the positive correlation of expected future short-term rates with inflation, which is mainly due to the highly positive correlation between the short-term nominal rate and inflation, the positive coefficient on the expectations-related component is, however, surprising. In fact, one explanation for a limited role of the expectations-related component for future real activity delivered by Favero et al. [17] is that the high positive correlation of the expected short-term rate with inflation implies a negative correlation of the expected nominal short-term rate with the real short-term rate due to the fact that the real short-term rate by construction has a negative correlation with inflation.

In fact, we find a positive relationship between expected future monetary policy and inflation for some of our countries. Figure 5 depicts the expected future interest rate for the Czech Republic, calculated as the actual short-term rate plus the summed up weighted expected short rate changes over the lifetime of the five-year bond. The correlation of the expected short-term rate with

Table 5: Pooled regression of future output growth

variable	coefficient	t-stat	Wald p-val.
$3m.\ yield$	-0.0732	-2.77	0.000046
$inflation$	0.0026	7.66	0.000000
TP	0.0088	2.43	0.000281
ER	0.2566	4.96	0.000397
Panel restr.			0.000000
Fixed Effects			
BL	1.6345		
CZ	5.5108		
HN	5.8671		
RM	14.8043		

$\Delta y_{t,t+12} = \alpha_0 + \alpha_1 i_t^{3m} + \alpha_2 \pi_t + \alpha_3 ER_t^{CS} + \alpha_4 TP_t^{CS} + u_t$

Sample: Mar'96 - Sep'04

R-squared 0.33, D.W. 0.41

inflation for the common sample between 1995 and 2004 is 47 percent for the Czech Republic. For Hungary it is 90 percent, for Romania 60 percent and for Bulgaria 49 percent.

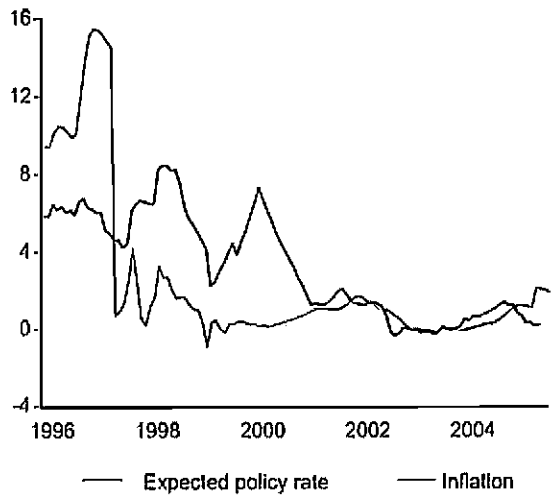
In our sample of countries the negative correlation between the expected future short-term rate and the real short-term rate does, however, not seem to nullify the predictive content of the expectations-related component of the yield spread for future real activity, although it has to be noted that the magnitude of GDP growth predicted by the expectations-related component is very small. An expectations-related component of 1% predicts only a GDP increase of roughly 25 basis points.

A more general problem might be a regime shift from a transitional to a more normal phase during our sample from 1995 to 2004. A rather tricky question is, when exactly the regime shift could have taken place. We tried with arbitrarily set dummies for a regime shift but these were never significant.

5.4.2 Term Premia of Twelve-month Interbank Rates

In some studies on the CEE countries' economies, interbank interest rates are used instead of government bond yields because they provide much more observations and offer longer time series than government bond or treasury bill data.

Figure 5: Forecast three-month treasury bill rate and inflation for the Czech Republic



Due to the scarcity of treasury bill data sets of sufficient length, we want to profit from the relatively long time-series available for interbank interest rates. We have to drop Bulgaria for which no interbank rates are available but in turn this allows us to add Poland and Slovakia, for which we have interbank rate data but only very few bond yield data, to our sample.

We estimate the Campbell and Shiller [5] VAR model from section 5.4 on our full sample ranging from 1995 to 2005 and then calculate the expectations-related component of the yield spread and, as a residual, the term premium component. Later on, we will derive a slightly more realistic measure of the expectations and term premium components by initialising the VAR from 1995 to 1999 and forecasting short-term rate changes out-of-sample between 2000 and 2004. This will allow us to mimic the information set which is actually available to the agents in a more realistic manner. In contrast to the preceding analysis on government bond yield data, our VAR now consists of the yield spread between twelve-month and one-month interbank interest rates, the monthly change in the one-month interbank rate and the spread of the one-month rate with respect to the Frankfurt one-month interbank rate. The spread of CEE interbank rates to the Frankfurt interbank rate should control for the fact that CEE countries' central banks are using policy rates for the managed float of their exchange rates

Figure 6: Term premia of interbank ($12m - 1m$) yield spread

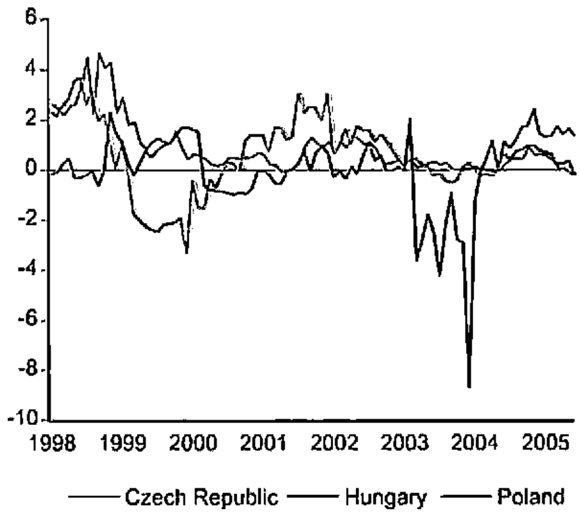


Table 6: Correlations between ER, TP and (1y-1m) yield spread

ρ	*CzechRep.*	*Hungary*	*Poland*	*Romania*	*Slovakia*
ER_t^{CS}, TP_t^{CS}	-0.74	-0.80	-0.90	-0.47	-0.81
$ER_t^{CS}, S_t^{(1y-1m)}$	0.14	0.08	0.50	0.37	-0.28
$TP_t^{CS}, S_t^{(1y-1m)}$	0.56	0.54	-0.08	0.66	0.79

S, ER and TP from full sample estimation 1995-2005

and, therefore, are up to a certain degree constraint by the conduct of monetary policy in eurozone and Germany, respectively.

Full sample estimation We estimate the VAR on the full sample from November 1995 through June 2005 for the Czech Republic, Hungary, Poland, Romania and Slovakia and then calculate the expectations-related component and, as a residual, the term premium component of the yield spread according to equation 14. We report the results for the term premia and the expectations-related components in figures 6 to 9 only from 1998 onwards because the volatility before this period is extremely high and would preclude the possibility of reading-off any relevant information from the graphs.

Figure 7: Term premia of interbank ($12m - 1m$) yield spread

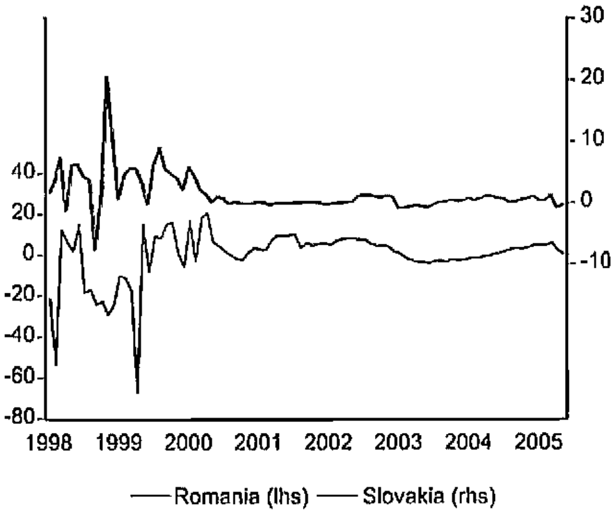


Figure 8: Expectations-related component of interbank ($12m - 1m$) yield spread

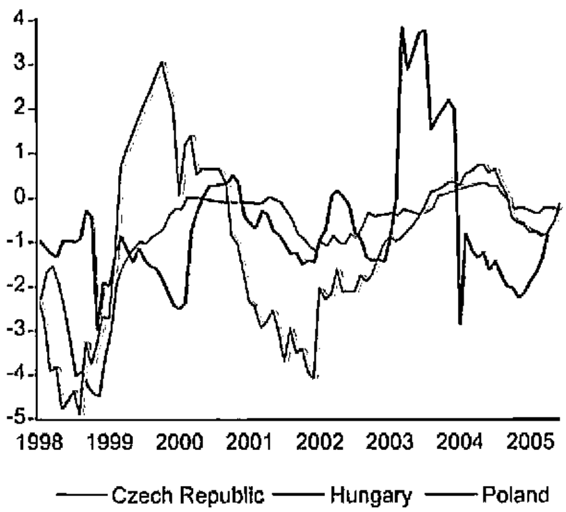
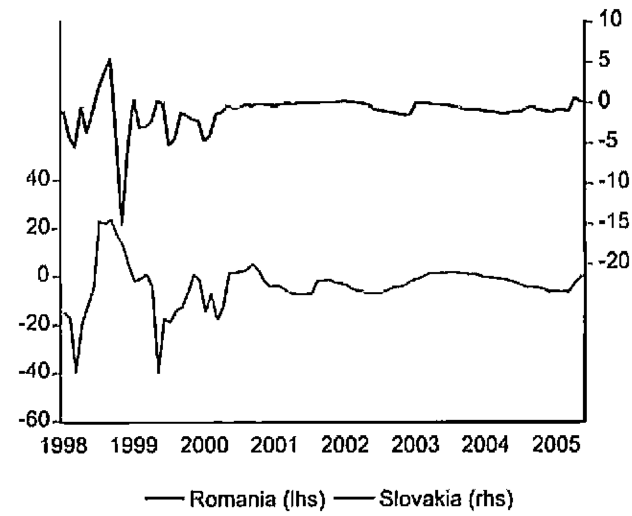


Figure 9: Expectations-related component of interbank $(12m - 1m)$ yield spread



As displayed by table 6 we can again extend a result of Favero et al. [17] to the CEE countries. The yield spread has a much higher correlation with the term premium component than with the expectations-related component, the only exception being Poland, where the term premium is not correlated with the $(1y - 1m)$ yield spread.

Returning once again to the seven forecasting equations of Favero et al. [17] from table 2, table 7 shows that in terms of the signs of the coefficients only the Czech Republic seems to be in the right ballpark. We abstract the equation containing the Hamilton Kim decomposition from the analysis of the interbank yield spread as its assumptions regarding the agents' information set when calculating the term premium are not adequate.

As can be seen from equation 6 in table 7, the ER component does play a role in determining future real GDP growth only in the Czech Republic and in Romania. Neither did the ER component play a role in Ang et al. [1] or in Favero et al. [17]. Favero et al. [17] raise the point that this could have to do with the lags of the transmission of monetary policy before feeding through to economic variables. They quickly dismiss this theory, however, as they do not find significant effects of the ER component on real GDP growth for horizons longer than one year in their own work.

The coefficient on the short-term interest rate instead, here the one-month interbank rate, has always the correct sign whereas inflation is never significant. This is not consistent with theoretical aggregate demand models with a short-term real interest rate because only the nominal part of the interest rate seems to be of importance here. More interesting, however, is the interpretation of the coefficient on the term premium in table 7. It is negative and significant for the Czech and the Slovak Republics and Romania but insignificant for Hungary and Poland. Thus, in contrast to the government bond yield model the term premium seems to predict output contractions rather than output expansions.

Regarding parameter stability, we do a Chow breakpoint test and a Chow forecast test. The two tests deliver, however, conflicting results. The breakpoint test rejects the null of no structural change in the middle of our sample in January 2000 for the Czech Republic, Romania and Poland. It does not reject the null for Slovakia and Hungary. The Chow forecast test for the sample January 2000 to September 2004, instead, rejects the null of no structural change for Poland and Romania only, however it does not reject the null for the Czech Republic. For the Czech Republic the results of the two tests are clearly conflicting. Therefore, in appendix A we also split the panel to analyse its stability over time.

Out-of-sample forecasts of the yield spread components In this section we estimate the VAR on the sample ranging from 1995 to 1999, then we forecast short-rate changes out-of-sample for the period 2000 to 2005 and calculate the expectations-related and the term premium components according to the same methodology as before.

Table 8 displays the correlations between the two components of the yield spread derived from the out-of-sample forecasts. Clearly, the connection between the yield spread and the term premium has become much stronger in the period between 2000 and 2005 than it has been in the second half of the 1990s, while the negative correlation between the ER and the TP components now has a lower magnitude than in table 6 referring to the estimation on the full sample. This could imply that the decomposition of the yield spread into an expectations-related and a term premium component, which should be orthogonal to each other, does not seem to work as well as in the above model.

Figures 10 to 12 depict the term premia and the expectations-related components derived from the out-of-sample forecasts of future short rate changes.

In the Czech Republic the term premium from the out-of-sample forecast

Table 7: GDP forecasting equations containing (1y-1m) interbank yield spread

		CZ	HN	PO	RM	SK
$eq1$	c	5.883^{***}	4.186^{***}	3.015^{***}	5.458^{***}	2.777^{***}
	S_t	0.777^{***}	0.006	-0.154	0.104^{*}	-0.034
	R^2	0.11	-0.01	-0.01	0.03	-0.01
$eq2$	c	5.481^{***}	4.977^{***}	2.923^{***}	17.275^{***}	6.599^{***}
	S_t	0.831^{***}	-0.011	-0.142	-0.459^{***}	-0.224^{**}
	i_t^{1m}	0.047	-0.053^{*}	0.006	-0.297^{***}	-0.333^{***}
	R^2	0.11	0.01	-0.02	0.36	0.22
$eq3$	c	5.585^{***}	4.842^{***}	3.095^{***}	16.908^{***}	6.505^{***}
	S_t	0.752^{***}	-0.008	-0.295	-0.460^{***}	-0.225^{**}
	i_t^{1m}	0.007	-0.016	-0.037	-0.232^{***}	-0.337^{***}
	π_t	0.545	-0.486^{*}	0.695^{**}	-0.978^{***}	0.214
	R^2	0.11	0.03	0.01	0.46	0.22
$eq4$	c	6.921^{***}	4.847^{***}	2.882^{***}	10.220^{***}	4.857^{***}
	i_t^{1m}	-0.173^{**}	-0.016	-0.007	-0.049^{*}	-0.154^{***}
	π_t	1.037^{**}	-0.486^{*}	0.600^{*}	-0.973^{***}	0.524
	R^2	0.04	0.04	0.02	0.29	0.05
$eq5$	c	5.880^{***}	4.834^{***}	3.131^{***}	13.881^{***}	6.426^{***}
	i_t^{1y}	-0.043	-0.016	-0.029	-0.153^{***}	-0.320^{***}
	π_t	0.886^{*}	-0.487^{*}	0.621^{*}	-0.713^{***}	0.204
	R^2	0.01	0.04	0.01	0.37	0.23
$eq6$	c	6.620^{***}	5.420^{***}	3.094^{***}	17.572^{***}	6.066^{***}
	i_t^{1m}	-0.410^{***}	-0.083^{**}	-0.037	-0.263^{***}	-0.311^{***}
	π_t	1.095^{**}	-0.455^{*}	0.690^{*}	-0.829^{***}	0.0923
	TP_t^{CS}	-0.536^{***}	0.018	-0.305	-0.468^{***}	-0.423^{***}
	ER_t^{CS}	-0.462^{***}	-0.504^{*}	-0.296	-0.570^{***}	-0.134
	R^2	0.40	0.12	-0.00	0.47	0.33
$eq7$	c	3.890^{***}	4.019^{***}	3.030^{***}	5.369^{***}	2.567^{***}
	TP_t^{CS}	0.069	0.072	-0.218	0.117^{*}	-0.234^{*}
	ER_t^{CS}	0.209	-0.209	-0.164	0.072	0.188
	R^2	0.00	0.02	-0.01	0.02	0.13
obs		107	107	107	93	107

R-squared reported is adjusted R-squared

*, **, and *** indicate significance at the 10%, 5%, 1% level respectively

Full specifiaction of equations is displayed in table 2

Table 8: Correlations between ER, TP and (1y-1m) interbank yield spread

ρ	*Czech.Rep.*	*Hungary*	*Poland*	*Romania*	*Slovakia*
ER_t^{CS}, TP_t^{CS}	-0.60	-0.15	-0.31	-0.25	-0.06
$ER_t^{CS}, S_t^{(1y-1m)}$	-0.25	-0.05	0.29	0.18	0.50
$TP_t^{CS}, S_t^{(1y-1m)}$	0.92	0.99	0.82	0.91	0.83

S, ER and TP from out-of-sample estimation 2000-2005

Figure 10: Term premia of interbank $(12m - 1m)$ yield spread

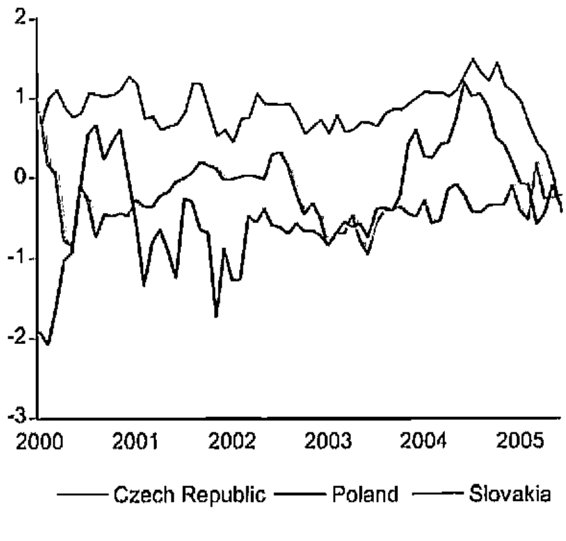


Figure 11: Term premia of interbank ($12m - 1m$) yield spread

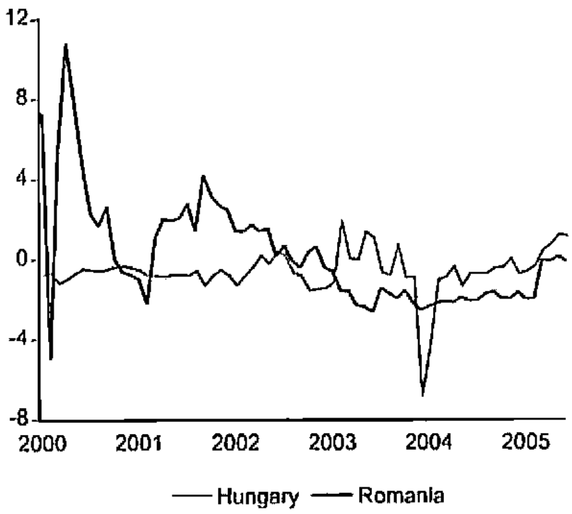


Figure 12: Expectations-related component of interbank ($12m - 1m$) yield spread

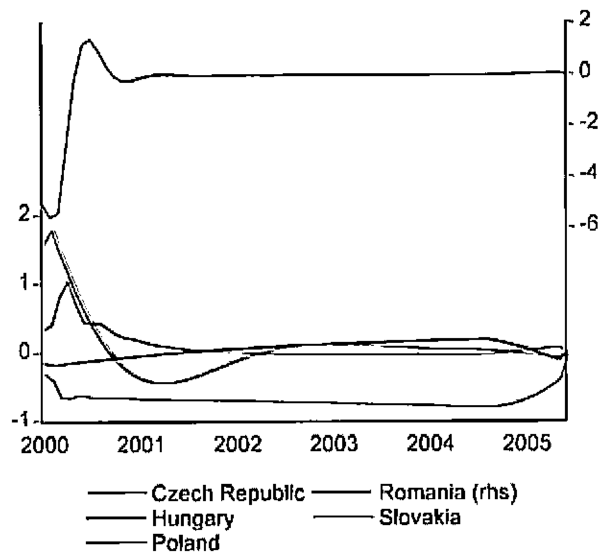


Table 9: Mean of TP and ER components between 2000 and 2005

	estimation	*CZ*	*HN*	*PO*	*RM*	*SK*
TP	*full sample*	0.42	−0.17	0.67	3.14	0.42
	out-of-sample	0.85	−0.61	−0.34	0.23	−0.27
ER	*full sample*	−0.27	−0.39	−0.94	−3.04	−0.58
	out-of-sample	−0.69	0.05	0.07	−0.26	0.09

'full sample' reports mean of respective yield spread component derived from VAR on full sample (Nov'95-Jun'05), while 'out-of--sample' reports mean of respective yield spread component derived on basis of VAR estimated from Nov'95 to Dec'99 and short-rate forecast out of sample from Jan'00 to Jun'05

model is higher than the term premia from the estimation on the full sample in section 5.4.2. On the other hand, for all other countries the term premia from the out-of-sample forecasts are lower than the term premia from the full sample forecasts. However, the results, which are displayed in table 9, probably are not due to the different methodologies used but due to generally lower term premia of the accession countries during the years 2000 to 2005 compared with the period from 1995 to 1999.

The big differences in term premia and expectations-related components in the two models imply as well differences for the predictive content of the yield spread's components, too. We will tackle these differences, however, only in the framework of a pooled regression in the next section.

Apart from the model estimated on the full sample in section 5.4.2 and the model estimated only on the first half of the sample in this section we did as well a VAR with a rolling estimation window starting with a sample ranging from November 1995 to December 1999 and then shifting the sample month-by-month to November 2000 through December 2004. This rolling VAR allows us to update the agents' information set with actual economic conditions on a monthly basis and to forecast short rate changes out-of sample and then calculate the TP and ER components. The results of the VAR with a rolling estimation window are, however, very similar to the results of the VAR estimated on the full sample in section 5.4.2. In fact, the TP and ER components of the rolling VAR are highly correlated (in a range of 80-90%) with the TP and ER components of the VAR on the full sample. Therefore, we do not report the results of the VAR with a rolling estimation window separately here.

Pooled regression A pooled regression of future GDP growth on the yield spread components, the one-month interbank rate and inflation naturally recalls the caveats on the use of panel data techniques alluded to briefly in section 5.4.1. The Wald test in tables 10 and 11 again rejects the identity restrictions on the coefficients and the panel restriction. Although the five countries are very different from each other we think that the pooled regression is still valuable because it puts us into a position to analyse the predictive content in the 'average' accession country.

In the regression of real GDP growth on the one-month interbank rate, the ER component and the TP component of the yield spread and the inflation rate in table 10, which is based on the estimation of the VAR on the full sample, the ER and the TP components suggest a negative relationship between the yield spread components and future economic activity. This casts some doubt on the information content of the 1y-1m interbank spread for future economic activity. Probably the one-year interbank rate is too short-term in order to contain significant information on agents' expectations on future output growth. The one-year rate probably is a too short maturity for a rate that should incorporate expectations on the future dynamics of the economy. A higher short-term nominal rate, correctly, implies lower future output growth and a higher inflation rate implies lower future output growth, too.

The results of the pooled regression are in contrast to our results from the model of the government bond yield and are very different to the results of Favero et. al. [17], Hamilton and Kim [21] and Ang et. al. [1], too.

While in the model of the government bond yield spread increases of the TP and ER components predict future economic expansions, now the TP and ER components predict future contractions of economic activity in the average CEE country. The sign of the coefficient on inflation has changed, too, from positive to negative, which is counter-intuitive as higher inflation and higher inflation premia inherent to long-term interest rates usually are associated with future output expansion and not with future output contraction.

In the second pooled model, predicting the interbank yield spread's components and the change in the one-month interbank rate out-of-sample for the period from 2000 to 2005, the term premium and the expectations-related component both predict future GDP contractions instead of future GDP expansions. Especially, the coefficient on the expectations-related component has a big magnitude as displayed in table 11. Inflation now comes in with a positive coefficient too. At the beginning of the new century inflation becomes an important pre-

Table 10: Full sample: pooled regression of future output growth

variable	coeff	t-stat	Wald p-val
$1m\ yield$	-0.2025	-8.21	0.000000
$inflation$	-0.0335	-2.06	0.000696
TP	-0.2857	-6.30	0.037101
ER	-0.3722	-6.16	0.077382
Panel restriction			0.000000
Fixed Effects			
CZ	5.3959		
HN	7.3704		
PO	6.3207		
RM	14.8500		
SK	5.0112		

Sample: Nov'95-Sep'04
R-squared 0.36, D.W. 0.48

Table 11: Out-of-sample: pooled regression of future output growth

variable	coeff	t-stat	Wald p-val
$1m\ yield$	-0.4432	-11.13	0.000133
$inflation$	0.2088	4.40	0.295978
TP	-0.5676	-4.31	0.000000
ER	-1.4134	-4.68	0.058426
Panel restriction			0.000000
Fixed Effects			
CZ	6.4519		
HN	7.2212		
PO	6.4469		
RM	17.8437		
SK	5.0692		

Sample: VAR Dec'95-Dec'99, forecast: Jan'00-Sep'04
R-squared 0.51, D.W. 0.83

dictor of future real economic activity in the average accession country, as it has been already for some time in the US and in other more developed economies. The positive coefficient on inflation could be due to less hawkish inflation targeting in the period between 2000 and 2005 compared to the second half of the 1990s. Whereas, in the 1990s inflationary pressures might have led to very restrictive monetary policy and, thus, were associated with future output contractions this does not seem to be any more the case for the beginning of the 2000s.

The fit of the out-of-sample model is largely superior to the fit of the model estimated on the full sample. This is thanks to the higher precision with which the coefficients on the short-term rate and inflation are measured.

Again, a problem which could apply here is parameter instability because countries first went through a transitional phase before entering in a more normal phase. Therefore, we put dummies for an arbitrarily chosen transition phase lasting from November 1995 through December 1999. These dummies are, however, never significant in the panel. Instead of Chow breakpoint and forecast tests which have shown to be problematic we simply break the panel in the middle in appendix A to analyse its stability.

6 Predicting Interbank Rates With a Taylor Rule

The predictive content of the yield spread has been shown to depend on the specification of the monetary policy reaction function, see e.g. Feroli [18]. We have found only a limited role for the predictive content of the yield spread in the five CEE countries analysed. In a panel, the term premium and the expectations-related component of the yield spread indicate future output contractions instead of future output expansions as one might have expected. This, of course, might be due to the specification of the monetary policy reaction functions in the accession countries under analysis.

We estimated the expectations-related component of the 1y-1m interbank yield spread for five countries using a VAR approach. Therefore, we could now extract expected interest rates from VAR forecasts of future one-month interbank rates. Small VAR models including only a short-term rate, the term spread and, where appropriate, the yield spread versus a benchmark rate are, however, known to have poor forecasting abilities. Their mixed results stem from the fact that they neither include the output gap, nor the deviation of inflation from target in the agents' information set. Therefore, we do not know how well the VAR predicts future policy rates.

In this respect, it is interesting to make use of a Taylor rule to check whether it delivers a satisfying description of the conduct of monetary policy in the five CEE countries. The Taylor rule provides us with an additional and, above all, a more precise measure of expected future monetary policy than the VAR. The predicted rates from the Taylor rule can then be compared to the correspondent expectations-related components extracted from a VAR forecast in order to compare the forecasting performance of a Taylor rule and a VAR decomposition regarding short-term interest rate forecasts.

Thus, we have to ask ourselves what the implicit reaction functions of the accession countries look like and how they compare to Taylor rules. We estimate a forward-looking Taylor rule on monthly data for the period between January 1998 and December 2002. We include the twelve-month ahead CPI inflation rate as a proxy for the central banks inflation expectations in our Taylor rule equation 15.

$$i_t = \rho r_{t-1} + (1-\rho)(i^* + \alpha(E_t\{\pi_{t,k}\} - \pi^*) + \beta y) + \varepsilon_t \tag{15}$$

Our preferred measure for modelling policy rates is the one-month interbank rate in each country. While it is true that this could cause some distortion in the sense that it is not a risk-free rate and that it varies according to the systemic risk inherent to the banking systems, treasury bill rates are not an option because their time series often start only in 2002 or 2003.

The coefficients on the inflation gap displayed in table 12 partly reflect the aggressive anti-inflationary policy usually observed in the first years after the switch to an inflation targeting strategy, which is due both to credibility problems and the fact that many central banks use inflation targeting for disinflation purposes.

The coefficients on the output gap, instead, are never significant between 1998 and 2002. This could reflect the fact that the five central banks were occupied with meeting inflation targets set in line with their countries' convergence processes and had only few leeway to stimulate output growth. Apart from representing central bank preferences, the estimated parameters represent, however, structural characteristics of the economy, too.

The results of the Taylor rule suggest that the yield spread between one-year and twelve-month interbank funds has only a limited capability for predicting

Table 12: Taylor rules of one-month interbank rates

	CZ	SK	PO	HN	RM
ρ	0.92***	0.93***	0.97***	0.97***	0.60***
α	1.35***	0.80**	2.38*	1.22**	0.76***
β	−0.03	0.76	−1.1	0.10	−0.18
R^2	0.99	0.62	0.96	0.97	0.71
DW	0.47	1.93	1.77	2.09	1.82

Estimation by GMM on the sample Jan'98 - Dec'02.
*, **, and *** indicate significance at the 10%, 5%and 1% level respectively. Twelve-month ahead inflation rate is instrumented by lags 1 and 12 of the output gap, the one-month interbank rate and its spread to DM/Euro rates, CPI inflation, the IMF commodity price index and the EUR and USD exchange rates.

future output growth but it should be highly relevant for future inflation dynamics instead.

In order to see whether a Taylor rule gives a good idea for the policy setting process in these countries we forecast policy rates from 2003 onwards. Based on the forward-looking Taylor rule of equation 15 which we estimate on the sample January 1998 to December 2002, we predict the one-month interbank rates in the Czech Republic, Poland, Hungary, Romania and Slovakia from January 2003 through September 2004. Our predictions are reported in figures 13 to 17.

In the Czech Republic the monetary tightening by the Czech National Bank due to accelerating investment demand, reviving exports, surging inflation, a relatively weak crown and buoyant domestic demand made the one-month PRIBOR rate increase at a stronger pace than predicted by our Taylor rule in 2004.

For Poland, as displayed in figure 14, the model does capture the Polish monetary easing in the first half of 2003 only with a quite long lag but the medium-term forecasting capabilities are satisfying.

Our forecast of the one-month interbank rate for Hungary, instead, performs badly compared to the forecasts of the other countries. As can be seen in figure 15, the Hungarian one-month interbank rate breaks the confidence bands of our forecast on several occasions. By all standards, 2003 was a year in which the Hungarian economy was characterised by high uncertainty regarding its medium-term economic prospects. A speculative attack on the Hungarian Forint in January, postponement of the EMU entry date and inflation way-

Figure 13: Predicted one-month interbank rate for the Czech Republic

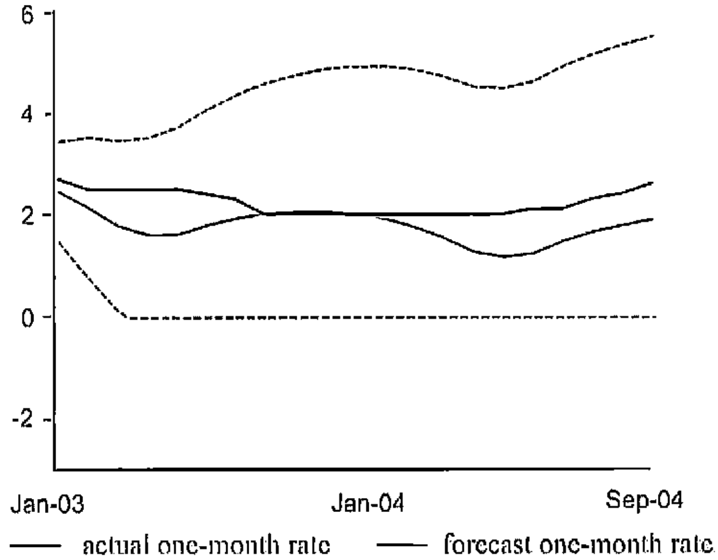


Figure 14: Predicted one-month interbank rate for Poland

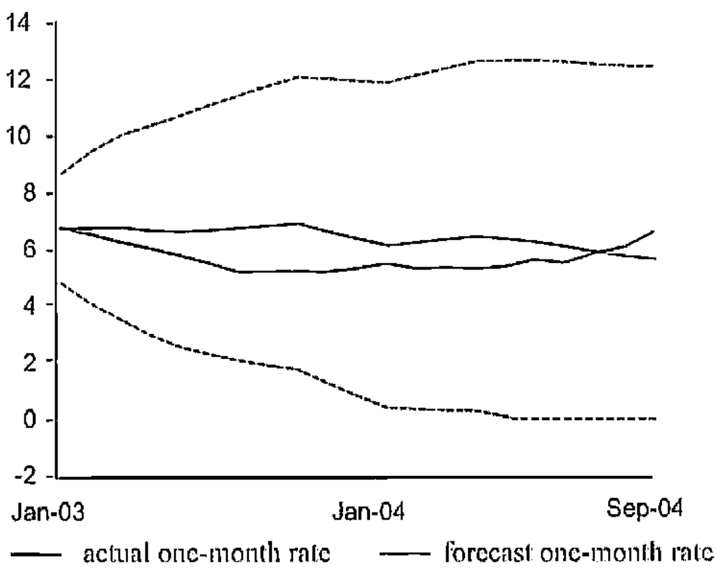
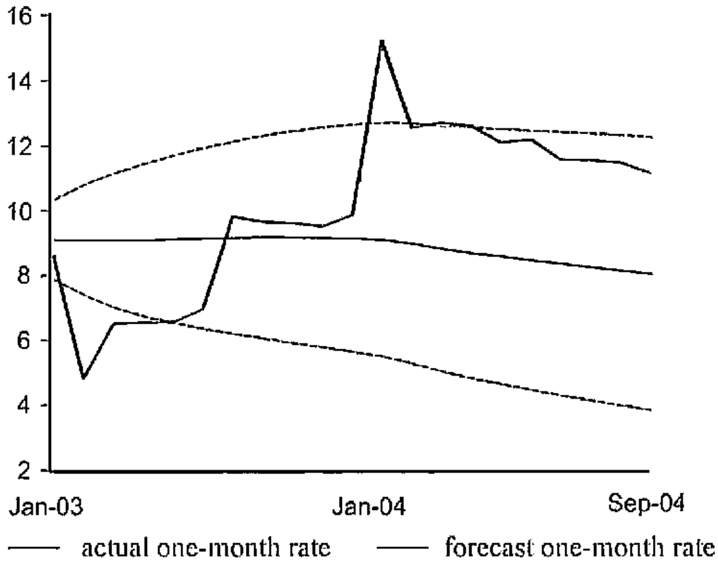


Figure 15: Predicted one-month interbank rate for Hungary



above market expectations caused financial turmoil which manifested itself in very strong interest rate volatility. In the aftermath of the speculative attack the one-month interbank rate jumped from 8.61% in January 2003 to 4.82% in February 2003 and then, later in the year, it bounced from 9.86% in November 2003 up to 15.19% in December 2003. Clearly, Hungarian policy rates were not very well described by a Taylor rule in 2003[3].

As can be seen from figure 16, the forecast one-month BUBOR interbank rate in Romania remains stubbornly high due to its historically high level of over 55 percent on average in the period between 1998 and 2002. Our forecast does not perform particularly well but remains, however, well inside the confidence bands.

The high inflationary pressures in 2003/04 push up our forecast rate for the Slovak Republic. At the end of 2004 the Slovak central bank succeeded to put inflation below four percent and lowered the policy rate on several occasions,

[3] This, however, cannot be said regarding an application of the Taylor rule for Hungary in general. As appendix B points out, estimation of the Taylor rule on the sample January 1998 - December 2003 and forecasting the one-month interbank rate for the year 2004, the forecast remains well inside the confidence bands. The incapacity of the Taylor rule to forecast 2003 rates is mainly due to the tumultuous Hungarian macroeconomic environment in 2003.

Figure 16: Predicted one-month interest rate for Romania

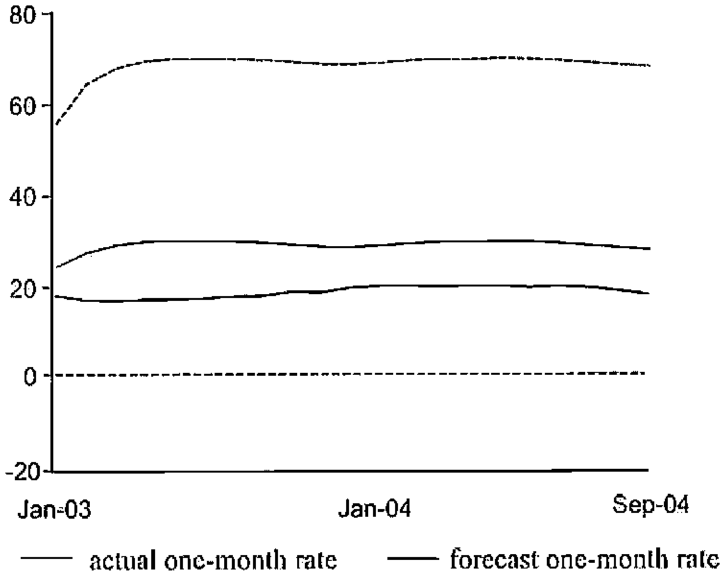
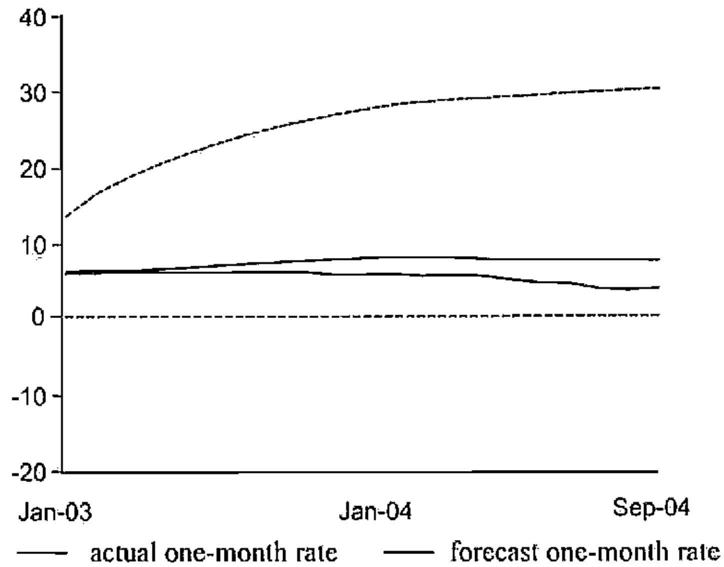


Figure 17: Predicted one-month interbank rate for Slovakia



too. Our model, incorporating rather high inflation expectations, is caught in the cold by the central bank easing and overestimates the BRIBOR one-month interbank rate by roughly four percentage points as can be seen from figure 17.

A backward-looking Taylor rule in which the monetary policy reaction function of the central bank includes the change in CPI inflation in the preceding twelve months delivers qualitatively similar results to the ones shown in figures 13 to 17 for the five countries.

Overall, the Taylor rules seem to describe the policy-setting process in the five CEE countries reasonably well. The central banks' emphasis probably has been mainly on inflation targeting and less on output stabilisation. The predictive content of the yield spread, however, has been shown to depend on the specification of the monetary policy reaction function. Due to the great emphasis given to disinflationary monetary policy, the yield spread in the five CEE countries retains probably only limited predictive content for future output growth.

Even if it would seem that the accession countries apply Taylor rules as rough guidelines for monetary policy, it needs to be asked how well the estimated parameters of the Taylor rules represent central bank preferences. Favero and Rovelli [16] raise the question, in as much estimated coefficients of monetary policy rules represent convolutions of central bank preference parameters, the structure of the economy and the efficiency in the conduct of monetary policy. Their discussion fully applies here. The estimated parameters of the Taylor rule might give a blurred picture of the deep central bank preferences if they represent convolutions of central bank, preferences, the structure of the shocks to the economy and the efficiency in the conduct of monetary policy. However, the present evidence suggests that the accession countries' one-month interbank rates and the interbank spread contain information on fututre output but above all information on future inflation.

7 Conclusions

We investigate the predictive ability of the yield spread and its expectations-related and term premium components for future real economic activity in five accession countries. To this end, we analyse the implications, which the components of the yield spread between five-year government bonds and three-month treasury bills and, alternatively, the components of the yield spread of one-year interbank rates over one-month interbank rates, have for the dynamics of future GDP growth.

Due to limited availability of long time-series data it is tricky to draw conclusions on the predictive content of the yield spread on the country level. It would seem, however, that the government term spread model suggests that the single components of the spread predict economic expansions in Bulgaria and in the Czech Republic, while they predict economic contractions in Hungary. The interbank term spread model, instead, appears to advocate that the single components of the spread predict economic contractions in the Czech Republic, in Hungary, in Romania and in Slovakia. The predictive content of the yield spread depends pretty much on the maturities chosen to construct the yield spread, on the country and as well on the period chosen to carry out the analysis.

A pooled regression for the model of the government bond yield spread, instead, suggests, that both the term premium component and the expectations-related component predict faster future real GDP growth, whereas in the model of the interbank yield spread both components are associated with slower future real GDP growth.

The pattern which emerges from both, analysis on the country level and a panel estimation approach is that the government bond yield spread indicates future expansions while the interbank yield spread indicates future contractions. While the results of the government bond yield spread model seem to go down well with earlier evidence found by similar studies on other countries, the finding of the interbank spread model might be more surprising. The question is whether the information content of the interbank spread for future economic activity is too limited for the purpose of predicting future real activity. This could be due to the fact that the one-year interest rate is essentially a short-term interest rate and, therefore, does not fully incorporate agents' expectations about medium and long-term developments of the economy and of future output growth.

A caveat to comparing the results of the two models is, however, that the models are based on two different cross-sections. While the model based on the government bond yield spread analyses data on the Czech Republic, Hungary, Romania and Bulgaria, the model of the interbank yield spread uses date from the Czech Republic, Hungary, Romania, Poland and Slovakia. In this context it would be of prime importance to check whether the differing results displayed by the pooled regressions of the interbank rate model and the government bond yield model are due to differing risk characteristics of banks and governments or whether it is rather an issue of the different maturities studied. Unfortunately, doing a counterfactual is not viable due to the lack of comparable data.

Having established rather mixed evidence on the predictive power of the yield spread, we employ a Taylor rule in order to reveal possible implications of the specification of the monetary policy reaction function for the predictive content of the yield spread. Predicting interest rates with Taylor rules it would seem that anti-inflationary policy is high up on the agenda while output stabilisation is only a secondary goal of the monetary authorities of the CEE countries analysed. Therefore, the yield spread might lose some of its predictive power for future real economic activity and probably would be more powerful for inflation forecasting.

A potential limitation to our analysis of the predictive content of the yield spread might be that simple forecasting relationships are typically quite unstable, presumably because the countries' economic structures, the conduct of policy, and the mix of shocks that buffet the economy can change over time. This seems to be particularly true for our choice of countries.

Further research on the agenda should take care of the fact that most of the work on the predictive content of the term spread for future inflation and output has been concentrated on the US which is known to be a relatively closed economy. Accession countries, however, are relatively small and open economies so that it would be interesting and necessary to study how the exchange-rate transmission channel, which makes open economies adjust comparatively rapidly to exogenous shocks, influences the results on the information content of the yield spread.

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Appendix A: Panel Split Analysis

We split the sample in two, fairly arbitrarily chosen, subsamples with a nearly identical number of panel observations. The first sub-sample ranges from November 1995 through December 1999 and the second from January 2000 through September 2004. For both sub-samples the coefficients on both the term premium component and the expectations-related component are negative and display similar coefficients and t-statistics, thus, indicating a certain stability in the relation between the two components of the yield spread and future economic activity over the whole sample from 19995 to 2004.

In contrast to the coefficients on the two components of the yield spread, the coefficient on inflation is instead negative in the period from November 1995 through December 1992, while it then turns around to become positive from January 2000 through September 2004. This is probably due to higher sensibility of the monetary authorities towards output stabilisation during the second half of our sample but could also partly be due to other changes in the struture of the economy apart from changes in central bank preferences. While in the first half of the sample higher inflation probably triggered market expectations of nominal rate hikes by the hands of a hawkish and anti-inflationary central bank and, subsequently, lower output growth, higher inflation in the second part of the sample might be associated with increased future economic activity supported by a more permissive monetary policy stance.

The adjusted R-squared in the second subsample is roughly 48%, up from 29% in the first sample, indicating a better model fit in the second subsample.

Based on this evidence, in particular regarding the development of the coefficient on inflation over time, we refrain from carrying out a more formal and

Table 13: Full sample: pooled regression of future output growth

	Nov'95-Dec'99		Jan'00-Sep'04	
variable	coeff	t-stat	coeff	t-stat
lm yield	−0.084	−2.45	−0.466	−10.927
inflation	−0.040	−2.05	0.219	4.474
TP	−0.166	−3.11	−0.618	−4.621
ER	−0.266	−3.69	−0.792	−5.754
Fixed Effects				
CZ	2.887		7.089	
HN	6.038		7.266	
PO	5.956		6.397	
RM	4.566		18.020	
SK	2.882		4.954	
	adj. R-squared 0.27		adj. R-squared 0.48	
	D.W. 0.29		D.W. 0.81	
	panel obs.: 236		panel obs.: 285	

thorough analysis to test parameter stability. It would seem clear that more formal tests should reject the hypothesis of non-time varying coefficients. In this context probably nothing else can be done about solving the problem of parameter instability than bearing in mind that the relations between the variables could be unstable over time.

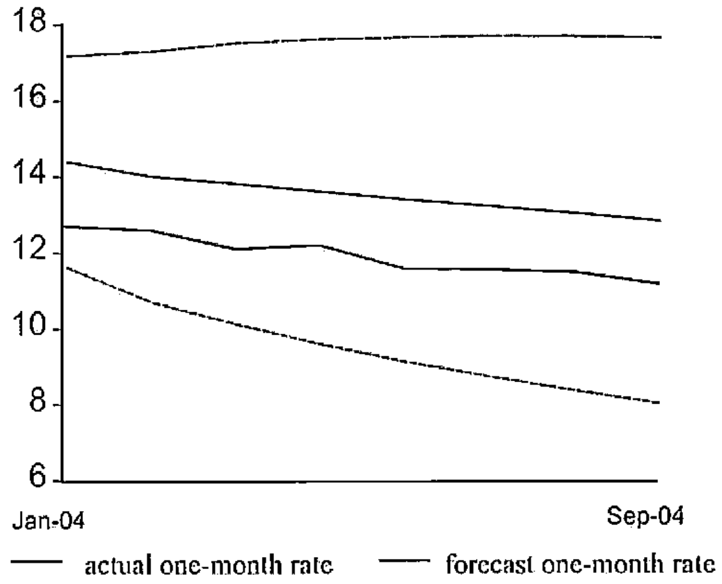
Appendix B: Hungarian Crisis 2003 and Prediciton of Interbank Rates

2003 was a year in which the Hungarian economy was characterised by high uncertainty regarding its medium-term economic prospects. A speculative attack on the Hungarian Forint in January, postponement of the EMU entry date and inflation way-above market expectations put financial markets into turmoil which manifested itself in very strong interest rate volatility. The worrying current account deficit squeezed investor confidence and raised the risk premium. This was accompanied by a scenario of adverse exchange rate expectations and steadily rising inflationary expectations throughout the year.

The distribution of the Reuters poll on the expected Hungarian EMU entry date[1] shows that financial markets' perception of the EMU entry date were

[1] The Reuters poll is available in the spreadsheets regarding the February 2004 Inflation Report of the Hungarian National Bank at http://english.mnb.hu/Resource.aspx?ResourceID=mnbfile&resourcename=inf_0402_i

Figure 18: Predicted one-month interbank rate for Hungary



shifted from 2007 to 2009/10 in the polls between January 2003 and January 2004 as can be seen from table 14. Increasing doubts on early EMU accession may have contributed to the upheaval in financial markets which showed up, for example, in the implied volatility of the Forint which peaked in January, June and December and in the extremely strong upward pressure on the yields of Hungarian government benchmarks in November and December 2003[5]. The postponement of the perceived EMU entry date contributed to adverse market reactions like pressure on the Hungarian Forint which, exacerbated by inflation well-above target, eventually led the Hungarian central bank to hike the base rate from 6.50% to 12.50% between 17 January and 28 November 2003 in order to restore its credibility.

To sum up, the high volatility of Hungarian interbank rates in 2003 and 2004, which resulted in more than trippled real rates during the year 2003, stemmed mainly from the central bank's efforts to stabilise the Forint in a narrow band accompanied by the likely postponement of eurozone accession from 2008 to

[5] The one-year benchmark yield, for example, was pushed from 9.90% on 24 November 2003 to 14.44% on 02 December

Table 14: Distribution of Hungarian EMU entry date as by Reuters polls

	2007	2008	2009	2010
Jan'03	50.0	42.3	7.7	0.0
Oct'03	0.0	14.3	85.7	0.0
Jan'04	0.0	0.0	80.0	20.0

reports percentage of respondents which indicated respective year as probable Hungarian EMU entry date

2009/10.

As displayed by figure 15, the Taylor rule did not perform well in this environment. Estimating the Taylor rule for the period January 1998 - December 2003 and forecasting the Hungarian one-month interbank rate for the year 2004 produces, however, forecasts which remain well-inside the confidence bands as can be seen from figure 18.