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FORECASTING REAL GDP: WHAT ROLE FOR NARROW MONEY?

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Contents

Abstract:

This paper analyses the information content of M1 for euro area real GDP since the beginning of the 1980s and reviews theoretical arguments on why real narrow money should help predict real GDP. We find that, unlike in the U.S., in the euro area, M1 has better and more robust forecasting properties for real GDP than yield spreads. This property persists when one controls for a number of other influences. We also evaluate the out-of-sample forecasting performance of different classes of VAR models comprising real M1, GDP and other indicators, using as benchmark a simple univariate model. As a result, only VARs in first differences are able to outperform the benchmark.

JEL: E41, E52, E58

Key words: Money; business cycle; forecast comparison; VAR models.

Non-technical summary

This study reviews the role of narrow money - predominantly real M1 - in terms of leading cyclical conditions in the euro area over the sample period 1981Q1 - 2001Q4. It also looks at how important real narrow money growth has been in relation to information already captured by short-term interest rates and a number of alternative leading indicator variables, such as various yield-spread measures, oil prices, stock price indices and exchange rates. Evidence is obtained on the basis of single-equation test regressions, as well as a more rigorous comparison of the out-of-sample forecast performance of different VAR-specifications (VARs in levels, first-difference VARs, VECMs and BVARs) with a univariate benchmark model (where GDP is forecast only on the basis of past information in GDP itself). These VARs comprise real M1 growth, GDP growth and the other potential leading indicator variables for real GDP.

As a result, the information content of different interest rate measures (long-term/short-term, real and nominal) for future changes in GDP seems rather mixed, while the predictive content of M1 remains important and tends to be largely unaltered in terms of coefficient size and forecasting horizon. Furthermore, in contrast to results obtained from studies for the U.S., in the euro area, M1 has better and more robust forecasting properties than the term spread (especially, when measured on the basis of historical euro-area interest rates). These properties of M1 are also maintained when comparing with a broader set of variables, like real and nominal effective exchange rates and oil prices. In all cases, narrow money seems crucial for cyclical developments. VARs in differences always outperform the benchmark irrespective of the forecast horizon considered, whereas the other classes of VAR models are not able to do so. Therefore, once the information content of money is taken into account, the choice of the general model seems to be more important than the selection of additional variables or the concrete forecast horizon. Furthermore, estimates of forecast models based on other variables, ignoring the information content of money, may be subject to a so-called "omitted-variable bias".

The paper also presents a number of theoretical arguments why money may affect cyclical conditions beyond the information already captured by interest rates. In this context it reconsiders conditions for the existence of a real balance effect. Furthermore it reviews the arguments recently re-stated by Meltzer (2001) and Nelson (2002a, 2002b) who focus on how money serves as a summary statistic reflecting a whole range of relative asset prices which are crucial for the monetary transmission process.

1. Introduction

Is money a good indicator of future real economic developments? And if yes, what are the theoretical justifications and what is the concrete mechanism behind the connection? These questions are among the most hotly debated in monetary economics. The present paper tries to shed some light on these issues for the euro area. In particular, it concentrates on the forecast performance of M1 for real GDP.

There are plenty of empirical papers dealing with this question for the US (see e.g. Hamilton and Kim (2002), Amato and Swanson (2001), Vilasuso (2000), Swanson (1998), Estrella and Mishkin (1997), Feldstein and Stock (1997), Friedman and Kuttner (1992)). The general conclusion is that M1 is not very useful in predicting future GDP growth[1] Compared with these studies, evidence on this with respect to euro area countries is scarce. For **Germany**, Kirchgässner and Savioz (2001) show that for four-quarter ahead forecasts of real GDP growth real M1 clearly outperforms forecasts based on interest rate spreads. This indicator role for M1 is also apparent in Sauer and Scheide (1995), who present evidence that there is a causal relationship from M1 to real economic activity measured by real domestic spending. Moreover, Fritsche and Kouzine (2002), find that M1 is one of the best leading indicators for business cycle turning points, measured by the index of industrial production, within a Markov switching model.[2] On the other hand, in the paper by Estrella and Mishkin (1997), the one-quarter growth of M1 is not significant in an equation forecasting the annualised growth rate of GDP four to eight quarters ahead. In Plosser and Rouwenhorst (1994), past and future monetary growth only helps to predict future growth in industrial production for relatively long forecasting horizons (five years). Furthermore, they show that the term spread is a significant predictor of future M1 growth. And finally, Seitz (1998), who looks at the best leading indicators for the growth of GDP

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[1] Exceptions are Swanson (1998), Vilasuso (2000) and Nelson (2002a). The first works with a special model selection procedure while the second uses detrended M1 growth that incorporates trend breaks. Nelson's paper is discussed in detail in section 2. Moreover, Leeper and Zha (2001), using a VAR analysis, conclude that the exclusion of money from this class of models is not empirically innocuous as the interpretation of the historical policy behaviour changes substantially once money is reintroduced. This result is confirmed by Favara and Giordani (2002), who examine the role played by shocks to the LM equation in shaping the dynamic behaviour of output, inflation and interest rates. A distinctive feature of their VAR analysis is that both the variables included in the system and the identifying restrictions used to isolate shocks to the LM equation are suggested by the class of models that assign a marginal role to monetary aggregates. They also find that all the variables they considered are not block-exogenous with respect to money. While Hamilton and Kim (2002) find that M1 is significant in explaining output growth in the U.S., this finding is not robust to all extensions of the information set.

[2] This is not true, however, when they use a probit model.

from the 1960s until the 1990s, shows that monetary aggregates do not play a significant role within a wide range of variables.

For **France**, Sauer and Scheide (1995), reveal a causal relationship between real M1 and real domestic spending within a cointegration framework whereas, in Estrella and Mishkin 1997, monetary aggregates are not helpful in predicting GDP irrespective of the chosen forecasting horizon. For **Italy**, Sauer and Scheide (1995), interpret evidence of a common trend in M1 and real economic activity as a special case of a causal role between the two variables. Furthermore, the interest rate spread does not contain any additional information on future output developments. Comparing the information content of the term spread and M1 for real GDP in Italy, Estrella and Mishkin (1997), reveal a slight puzzle in that the spread only becomes significant once M1 is added to the relation, although the monetary aggregate itself is mostly insignificant and has the wrong sign. Altissimo et al. (2002), use two approaches to analyse the relation between surprises in GDP and innovations in monetary variables. The first requires filtering the new information contained in monetary variables by mapping surprises into estimates of the structural disturbances impinging on the variables of interest and then starting a new forecasting round of the model; the second looks directly at the correlations among surprises. The monetary variables taken into account are M2 and the currency component of M1. Within the first approach neither M2 nor currency contribute to reducing the forecast uncertainty on GDP. In contrast, the second approach reveals that there is information in the two monetary aggregates for forecasting real GDP.

Finally, Canova and de Nicoló (2002), assess the importance of different monetary disturbances as sources of cyclical movements for the G-7 from 1973 to 1995. For that purpose they use a VAR model with industrial production as a proxy for real activity, real M1, the term spread and inflation. The major result of their paper is that the combined contribution of these monetary disturbances for real economic fluctuations is large in Germany and Italy. In Germany, there is a single monetary shock which explains the major part of output variability. In contrast, in France, monetary disturbances hardly contribute to output fluctuations. These conclusions are robust to the choice of sample period and to the inclusion of further variables, especially stock returns and short- and long-term nominal interest rates. The peculiarity with Canova and de Nicoló's approach is that monetary disturbances are an amalgam of many different factors, not just M1.

Overall, the results concerning the information content of M1 for real activity in general and real GDP in particular in euro area countries are not conclusive. Furthermore, up to now, there were only a few euro area countries under investigation.[3]

This study differs from the aforementioned ones in several respects. First, the role of narrow money for output has so far not been studied for the whole euro area. There are several papers dealing with the situation in individual euro area countries (see the discussion above) but the results may differ for the euro area as a whole. Second, we distinguish between different forecasting horizons ranging from one quarter to two years. Usually, only one such horizon is evaluated.[4] Third, in assessing the role of M1 for output we perform an ex-post and an ex-ante analysis. And fourth, we compare the forecasting ability of different optimal VARs because the time series properties of the data and the results from preliminary model analysis are ambiguous.

The paper is structured as follows. The next section presents an overview of theoretical arguments why money might be useful for forecasting real GDP, effects of monetary policy apart. The third section contains the empirical analysis. In this part, we first present some preliminary evidence that M1 might be useful for forecasting GDP by drawing on a recent study by Hamilton and Kim (2002). In a second step, we derive our univariate benchmark model against which we judge several VAR models in terms of their ability to predict GDP out of sample. The last chapter summarises and draws some tentative conclusions.

2. Why does money help forecast GDP – some theoretical arguments

In most modern macro models, short-term interest rates rather than monetary quantities capture the monetary policy stance. Monetary policy rules, for example, are typically specified in terms of a money market interest rate.[5] This raises the question of how money in general (and in particular M1, which is the aggregate we focus on) might be useful in explaining and predicting business fluctuations over and above the influence of interest rates. In what follows, we present some of these arguments emphasised in the literature.

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[3] There are also papers which try to construct composite leading business cycle indicators in which different measures of money enter, see e.g. Berk and Bikker (1995), for an analysis for, inter alia, several EU countries.

[4] Exceptions are Swanson (1998), for the US and Estrella and Mishkin (1997), as well as Plosser and Rouwenhorst (1994), in multi-country studies. Bagshaw (1985), compares the predictive performance of M1 relative to a univariate model of GNP and finds no significant differences. Furthermore, he stresses that the merits of M1 in helping to forecast output depends more on the forecasting period considered than on the forecast steps or the frequency over which M1 is measured (monthly versus quarterly).

2.1. The traditional real balance effect

Probably the first theoretical considerations about this relation originate from Pigou (1943) and Patinkin (1965). What has become known as the famous Pigou- or the real-balance effect, describes wealth effects created by a change in the stock of real money. Patinkin (1965), p. 20, defines the real balance effect as "an increase in the quantity of money, other things being held constant, [that] influences the demand for a commodity just like any other increase in wealth". There has been a long debate as to whether or not money should be treated as a part of wealth. Gurley and Shaw (1960), tried to clarify this discussion by introducing the distinction between outside and inside money.[6] Outside money is a part of government debt (including the central bank), while it is an asset of the private sector. An example is currency in circulation. Inside money constitutes debt of private agents as well as an asset held by them. This is true, for example, for overnight deposits. Ignoring distributional effects and the efficiency increasing effects of money compared to a barter economy, wealth effects would, if at all, only apply to outside money[7]. This would imply that we should not pay attention to monetary aggregates per se, but to its breakdown into inside and outside money. The latter is normally only a small part of total monetary aggregates; for example, in 2001, currency was only about 15 % of euro area M1 and only 2 % of total financial assets of the non-financial sector. It seems doubtful that such a small aggregate can have a significant impact on the economy. The judgement would be different, however, if both inside and outside money mattered.

2.2. Money in modern macro models

In all existing economies, private parties hold transaction balances despite the fact that they yield a lower return than other very short-term riskless assets. This indicates that there must be advantages from holding money which have so far not been considered in our discussion. The advantages relate to the assumption that money holdings facilitate transactions and lower

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[5] In the volume by Taylor (1999), all the papers presented use such a formulation.

[6] Whether these wealth effects are only a transitory phenomenon or are also relevant in the long run is crucially dependent on the assumptions of money being the only financial asset or not and on the time horizon considered for individual decisions. If bonds co-exist with money and agents have an infinite horizon the wealth effects may even be existent in the long run, see Handa (2000, p. 492).

[7] Furthermore, one has to recognise that households fail to take into account the impact that a future creation of excess money may have on the economy.

transaction costs.[8] This idea has been incorporated in money-in-the-utility-function models (see e.g. Woodford (2003, ch. 2)), shopping time models (see e.g. Bakhshi et al. (2002)) or cash-in-advance models with population growth (Ireland (2002)).[9] Within the first class of models, originating from Sidrauski (1967), the household maximises the expected value of the discounted sum of per-period contributions to utility u of the form

$$(1) \qquad E_0\left[\sum_{t=0}^{\infty} \beta^t u(c_t, m_t)\right],$$

where $0 < \beta < 1$ is the discount factor, E is the expectation operator and the per-period utility u depends positively on consumption c and real balances $m = M/P$, subject to its intertemporal budget constraint, possibly incorporating a borrowing limit (Woodford (2003, ch. 2)). The way money affects the consumption path crucially depends on the assumption made about u_{cm}.

If $u(.)$ is additively separable between its arguments c and m, the marginal utility of consumption would be independent of real balances, just as in the case of a cashless economy. This implies that aggregate demand and the expectational IS curve would be unaffected by real money balances. There would be no real balance effect despite the fact that money enters the utility function. One alternative way of justifying the neglect of real balance effects on the marginal utility of consumption is Woodford's (2003, ch. 2.3.4), case of a "cashless limiting economy". In this model the marginal utility of additional real balances becomes quite large as household real balances fall to zero, so that it is possible in equilibrium to have a non-trivial interest-rate differential between monetary and non-monetary assets. Yet, at the same time, the transactions that money is used for are sufficiently unimportant so that variations in the level of real balances have only a negligible effect on the marginal utility of consumption. The idea is that in such an economy money is used for transactions of only a very few kinds, though it is essential for those. As a result, positive real balances are demanded even in the case of a substantial interest-rate differential (and hence, a substantial opportunity cost of holding money); but equilibrium real balances are very small relative to national income. Consider an economy in which a fraction α of goods may only be purchased with cash and make the parameter α

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[8] For King (2002), even the proof of a significant role for money for real developments has to be based on the two observations that money reduces transactions costs and that transactions costs are important in determining asset prices. The frictions which money helps to overcome in financial markets have to be related to its role in providing liquidity services.

arbitrarily small. This means that whereas the elasticity of u_m with respect to real consumption (ε_c) is positive[10]

$$(2) \qquad \varepsilon_c = \frac{c \cdot u_{mc}}{u_m} > 0,$$

the elasticity of u_c with respect to real balances (ε_m)

$$(3) \qquad \varepsilon_m = \frac{m \cdot u_{cm}}{u_c},$$

which is essential for a real balance effect to be operative, is infinitesimally small. Woodford (1998), shows that if monetary policy may be characterised by an interest rate rule which specifies the short-term nominal interest rate as a function of the price level, there is no further need to consider the role of money in such an economy.

In the cashless-limit environment as well as in the consumption-money-separability environment money is redundant, as real money balances are demand determined given the endogenously determined interest rates and output. The usual LM equation in such a framework serves the sole purpose of determining the quantity of money the central bank needs to supply to clear the money market. The cashless-limiting-economy assumption, however, seems questionable as e.g. currency certainly provides valuable services to consumers. These may stem from its anonymity or from the fact that exchanges conducted with money can be done "without knowledge of individual histories" (Wallace (2000)). Against this background, McCallum (2000, 2001, 2002) strongly argues that there is also no compelling theoretical basis for the assumption of separability of u (see also Woodford (2003, ch. 2.3.4)). Thus, a direct money effect would arise

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[9] Croushore (1993), shows that the first two models are functionally equivalent. Holman (1998), postulates that money-in-the utility-function models allow for transactions as well as precautionary and store-of-value motives for holding money. McCallum (2000) presents a reduced form shorthand of all these analyses by introducing a transactions cost function, which reflects the transaction-facilitating properties of money, in the per-period budget constraint.

[10] Assuming separability would imply that the elasticity ε_e is negligible.

if real balances enter the representative agent's utility function, which in turn is not additively separable in consumption and real balances, but has a positive cross-derivative, i.e. $u_{cm} > 0$.[11]

This would result in an (expectational) IS function depending on the real rate of interest, expected future output, government expenditures as well as real balances.

2.3. Money as a proxy for a whole range of relative prices of assets

Friedman and Schwartz (1963), Meltzer (2001), and Nelson (2002a, 2002b) evaluate the important role of money for real activity and prices from a more general perspective. In particular, they argue that money cannot be ignored because it proxies the effects of many other asset prices on aggregate demand. If changes in money lead to changes in private sector portfolios and changes in yields of financial and real assets, this in turn also influences real spending decisions. The usefulness of this channel and the portfolio balance effect that arises depends on the assets being imperfect substitutes. A special advantage of this kind of model is that it tries to identify separate effects of real money on aggregate demand which are not captured by a short-term real interest rate.

What are the theoretical rationalisations for a direct money term in aggregate demand functions? Meltzer (2001), Friedman and Schwartz (1982), and Brunner and Meltzer (1993), state that money demand is not only dependent on one interest rate but a function of many different asset yields and wealth including human and non-human wealth. Therefore, money plays a special role for real developments in that it proxies the effects of these different yields and wealth effects which are relevant for economic activity.[12] In Meltzer's view, the gap between desired and actual real balances is a measure of the relative price adjustment necessary to restore the new full equilibrium. And, as he argues, a measure of the real money stock serves as a good summary statistic of the various changes in yields and wealth. The relevant yields include the whole term structure of interest rates, the yield on shares, the exchange rate, yields on housing etc.

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[11] The empirical analysis for the US in Koenig (1990), strongly suggests $u_{cm} > 0$. Taking the effects of real balances, measured by M1, into account, there is little evidence that other variables like anticipated or lagged changes in income, stock prices, government purchases or any other variable that might influence u_c or serve as a measure of liquidity, have a direct impact on consumption or its timing. In Koenig's study the effect of real balances on consumption is quite strong: a 10 percent increase in real M1 results in a three percent increase in spending on non-durables and services. This kind of non-separability has already been considered by Sidrauski (1967). Ireland (2002), however, derives a real balance effect in an infinite-horizon optimising model which does not require non-separability in utility.

A model which tries to fully capture the role of money therefore has to incorporate multiple assets which are imperfect substitutes for each other. If the yields influencing money demand and the yields influencing aggregate demand are correlated, real balances will be a good indicator of real developments in that they summarise all the relevant aspects.

Recent applications of this general idea may be found in Nelson (2002a, 2002b). Nelson (2002a) derives this effect within an intertemporal general equilibrium model with Calvo price setting where money becomes important in explaining output once portfolio adjustment costs are present. In Nelson (2002b), this is the case in an environment where current private sector shocks are not observable to the monetary authority. In such a world, looking at money is information-efficient and aids inflation stabilisation if current period nominal money growth can be observed.[13]

The special role of money may be yet more important when nominal interest rates are close to zero. Meltzer (2001), argues that a monetary expansion can stimulate the economy even in this case as nominal securities are not the only substitute for money. If, at some point, households and firms become satiated with money balances at the current level of income, any attempt to increase the money supply leads them to adjust their portfolios to limit their money holdings. These portfolio changes lead to changes in relative yields on financial and real assets and hence on real spending. The essential question then is whether there exists such a satiation point (King (2002)).[14]

3. Empirical analysis

In the following, we first present some single-equation evidence in the spirit of Hamilton and Kim (2002) that money may be useful in forecasting real GDP growth in the euro area. This part also enables us to compare evidence from the euro area and the US. The next section develops a univariate benchmark model against which we assess the forecast performance of a battery of VARs.

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[12] The results of Coenen et al. (2001) within a New-Keynesian passive-money-type approach can be interpreted in the light of this argument. In their model money is a helpful summary statistic for uncertain real output as money demand depends on output.

[13] In the same spirit, Tödter (2002) presents a model in which money is a summary statistic for a whole range of shocks hitting the economy.

3.1. Some preliminary evidence

Hamilton and Kim (2002), establish the importance of the yield spread for forecasting real output growth in the United States for the period 1953:Q2 to 1998:Q2. They use the following equation:

$$(7) \qquad \Delta y_t^h = \alpha_0 + \alpha_1 (il - is)_t + \alpha_2 x_t + \varepsilon_t \,,$$

where $\Delta y_t^h = \dfrac{400}{h} \left(\ln Y_{t+h} - \ln Y_t \right)$ is the annualised real GDP growth over the next h quarters and x_t is a vector of alternative explanatory variables (e.g. M1) besides the term spread (il - is). Their general conclusion is that the term spread is especially useful in predicting real GDP growth up to two years ahead.

In what follows we estimate one specific test regression for the euro area. It reads as

$$(8) \qquad \Delta y_t^h = \alpha_0 + \alpha_1 spread_t + \alpha_2 \Delta_4 m1r_t + \alpha_3 x_t + \varepsilon_t \,,$$

where $m1r$ is seasonally adjusted real M1 (h = 1,...,8). Our GDP measure (y) refers to seasonally adjusted GDP (see Figures 1 and 2 below).[15] As regards the spread, we do not only assess the performance of the long-term (the yield on 10-years government bonds) minus the short-term rate (the 3-month money market rate) (il-is), but also the difference between a composite lending rate and the money market rate (clr-is), the difference between the capital market rate and the own rate of M1 (il-i_{m1}), the difference between the money market rate and the own rate of M1 (is-i_{m1}) as well as the difference between the German yield on bonds and the German money market

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[14] The dependence of the IS relation on money may also be rationalised within the credit channel framework as credit is the main balance sheet counterpart to money (Bernanke and Blinder (1988)). However, at times when money and credit exhibit disparate movements, it seems worthwhile to treat both separately.

[15] Historical data for the euro area over the period 1980 Q1-1998 Q4 are constructed on the basis of national series converted into euro using the irrevocably fixed exchange rates of 31 December 1998. For a discussion of the merits and a presentation of various aggregation methods for historical euro area data, see Brand, Gerdesmeier, Roffia (2002, pp. 28-29), and, in particular, Beyer, Doornik and Hendry (2000, 2001). The quarterly growth rates calculated from the euro-area-11 GDP series have been used to extend the euro-area-12 GDP observations from 2000 Q4 back to 1980 Q1. (Greek data have been converted on the basis of the irrevocable fixed exchange rate determined on 19 June 2000 for Greece). For more detailed information on the compilation of these data and, in particular, the underlying national data, see Brand, Gerdesmeier and Roffia (2002, Annex D.2, pp. 56-58) and Gerdesmeier and Roffia (2003). The monetary data have been aggregated in line with the method used in Brand, Gerdesmeier and Roffia (2002, Annex D.1, p.56). Euro-11 (euro-12 from January 2001) ten-year government bond yields are weighted averages of long-term interest rates using national shares in M3 as weights. (Source for national series Datastream and BIS, based on national 10-year government bond yield or closest substitutes. 3-month money market interest rates have been constructed in the same manner (Source: BIS), and linked to EURIBOR (from Reuters) as from 1999 onwards. The world-wide stock price index has been taken from Datastream. Source for the oil price index, the EUR/US-$ exchange rate, and German interest rates: BIS. These data are available from the authors upon request. Furthermore, the own rate of M1 has been constructed on the basis of Stracca (2001), the nominal and real effective exchange rate measures on the basis of Maeso-Fernandez, Osbat and Schnatz (2001), the HICP data on the basis of Gerdesmeier and Roffia (2003), the composite lending rate on the basis of Calza, Manrique, and Sousa (2003), and the Divisia aggregate on the basis of Stracca (2001).

rate (*il_ger-is_ger*).[16] Nominal interest rates have been divided by 400. *Ex-post* real interest rates are calculated on the basis of nominal interest rates and annual changes in the GDP deflator (*pgdp*) (or the HICP) (divided by 4). In one case, M1 is substituted by a Divisia aggregate (*divr*). Finally, x_t are additional variables like stock prices (*stock*) (a worldwide stock price index highly correlated with the Euro Stoxx which could not be used as it only exists since the end of the 80s), oil prices in US-Dollars (*oilp*) and in euros (*oilp€*), the bilateral US-$/€euro exchange rate (*e*) as well as the multilateral effective nominal (*eeffn*) and real (*eeffr*) exchange rates of the euro. All data, except interest rates, are in logarithms. Our sample runs from 1980:Q1 to 2001:Q4.

Figure 1: Real M1 and GDP: Log Levels
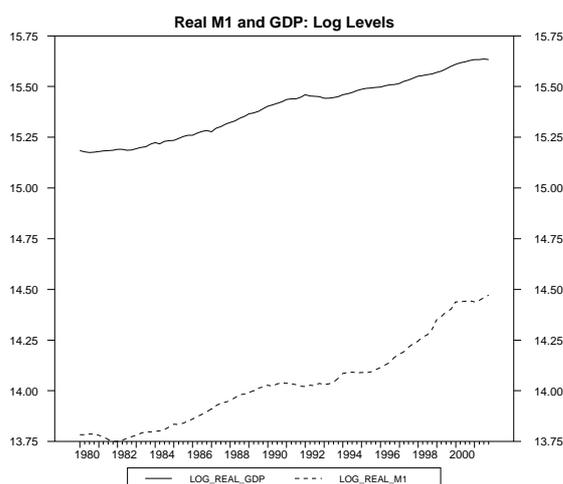
Figure 2: Real M1 and GDP: Changes on a Year Earlier

Real M1 and GDP: Log Levels

LOG_REAL_GDP — — — LOG_REAL_M1

Real M1 and GDP: Changes on a Year Earlier

R_GDP_GROWTH_YOY — — — R_M1_GROWTH_YOY

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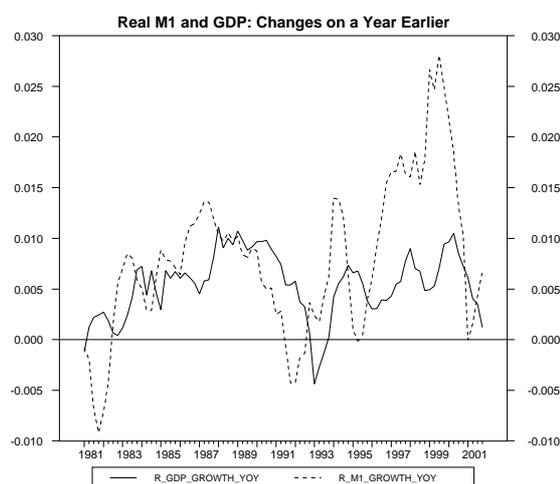
[16] The inclusion of the German spread may be rationalised by the fact that within the former (asymmetric) European Exchange Rate Mechanism the Bundesbank pursued an independent monetary policy aimed at price stability while the other ERM countries tried to maintain a stable exchange rate vis-à-vis the Deutsche Mark (De Grauwe (2000, ch. 5), Wellink and Knot (1996)). If Uncovered Interest Parity holds, it seems natural to consider the German spread.

Table 1 shows the coefficient estimates of equations (7) or (8), the (Newey-West-corrected) t-ratios and the adjusted R² of the equations. Each model block includes results for the forecast horizons $h = 1, …, 8$. For model 1, which only takes into account the term spread, the estimated coefficients are not significant at the 5 percent test level for all $h = 1, …, 8$. This is in contrast to the results of Hamilton and Kim (2002) for the US.[17] Adding the growth rates of real M1 to the test equations does not alter the results for the spread (model 2). However, the estimated coefficients of M1 are significant at the 1 percent test level for all $h = 1, …, 8$. The adjusted R² are considerably higher than those of the first model block.[18] The spread between the capital market rate and the own rate of M1 (model 3) is more important for $h = 1$ to 4. The real capital market rate based on the HICP (model 5) affects the growth rate of output significantly, whereas the real money market rate has no significant influence (model 6). All these approaches have one feature in common: in all cases the growth rate of M1 remains significant. This is also true if the regression equations only include M1 (model 4). This evidence does not change if the equations additionally contain other variables like the annual change of the world stock prices, of oil prices, of the US-\$/€ exchange rate, of the nominal or real effective exchange rates or of other interest rate variables. These other variables do not help to forecast output in the short term ($h = 1$ to 4). In the longer run ($h = 4$ to 8) some variables have significant coefficients, especially, the nominal effective exchange rate and the oil prices in euro. These results are also available upon request. It is also worth noting that the results for the change of the real Divisia aggregate (model 7) point in the same direction as the M1 change.

In sum, this exercise gives preliminary evidence of the importance of M1 for future output developments in the euro area. However, the forecast performance is not systematically evaluated relative to a benchmark. Moreover, the analysis is only carried out within a single equation approach. These issues are addressed in the following sections.

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[17] Evidence about the information content of various financial spread measures with respect to a large number of EU countries has been found in Davis and Fagan (1997). Their results are not supportive for the hypothesis that financial spreads can be used comprehensively and indiscriminately as indicators of future output growth and inflation. In the few cases where Granger-causality has been detected, the forecast performance proved to be disappointing, and, furthermore, their forecasting equations appeared to be unstable.

[18] On the other hand, the German interest rate spread helps to forecast changes in output over time for horizons $h = 6$ to 8, whereas the difference between the money market rate and the own rate of return of M1 (between a composite lending rate and the money market rate) does so for the horizon $h = 2, 3$ ($h = 1$). These results are available from the authors upon request.

Table 1: Coefficient estimates of equations (7) and (8) for future real GDP (Hamilton/Kim approach)

Model	Variables	Forecast horizon h							
		1	2	3	4	5	6	7	8
1	il-is	.462	.471	.451	.402	.362	.323	.293	.273
		(1.638)	(1.783)	(1.813)	(1.693)	(1.636)	(1.614)	(1.625)	(1.658)
	adj. R²	.050	.093	.115	.108	.097	.087	.083	.082
2	il-is	.118	.098	.077	.050	.043	.052	.058	.064
		(.425)	(.414)	(.374)	(.270)	(.260)	(.358)	(.443)	(.534)
	$\Delta_4\log(m1r)$	.260	.280	.279	.263	.239	.205	.178	.157
		(4.401)	(5.454)	(5.663)	(5.085)	(4.406)	(3.632)	(3.120)	(2.750)
	adj. R²	.165	.320	.418	.427	.391	.331	.293	.266
3	il-i$_{m1}$	.296	.283	.261	.208	.165	.097	.028	-.007
		(2.770)	(3.292)	(2.920)	(2.074)	(1.449)	(.752)	(.243)	(.063)
	$\Delta_4\log(m1r)$	.403	.414	.401	.360	.319	.258	.200	.165
		(5.569)	(6.126)	(5.535)	(4.446)	(3.554)	(2.549)	(1.977)	(1.668)
	adj. R²	.202	.376	.479	.471	.419	.339	.291	.262
4	$\Delta_4\log(m1r)$	.280	.297	.292	.271	.246	.214	.187	.168
		(5.038)	(5.575)	(5.510)	(4.876)	(4.25)	(3.544)	(3.072)	(2.771)
	adj. R²	.172	.325	.422	.433	.397	.337	.299	.272
5	il$_{real(HICP)}$	.441	.436	.437	.408	.421	.417	.377	.346
		(3.327)	(3.987)	(4.160)	(3.395)	(2.925)	(2.342)	(1.992)	(1.845)
	$\Delta_4\log(m1r)$	.363	.376	.361	.330	.310	.282	.246	.223
		(6.431)	(7.474)	(6.902)	(5.451)	(4.561)	(3.657)	(3.163)	(3.036)
	adj. R²	.215	.398	.500	.492	.458	.391	.332	.273
6	is$_{real(HICP)}$	.211	.217	.223	.219	.221	.195	.153	.125
		(1.038)	(1.217)	(1.367)	(1.453)	(1.521)	(1.394)	(1.159)	(.983)
	$\Delta_4\log(m1r)$	.359	.374	.361	.331	.308	.268	.225	.200
		(5.814)	(7.502)	(7.484)	(6.107)	(5.038)	(3.992)	(3.341)	(2.972)
	adj. R²	.178	.340	.430	.430	.395	.321	.265	.236
7	il-is	.219	.200	.177	.148	.134	.132	.132	.134
		(.835)	(.894)	(.918)	(.858)	(.870)	(.972)	(1.054)	(1.144)
	$\Delta_4\log(divr)$	.478	.531	.534	.495	.447	.380	.321	.275
		(4.761)	(6.108)	(6.276)	(5.494)	(4.645)	(3.729)	(3.100)	(2.669)
	adj. R²	.183	.372	.494	.496	.449	.373	.318	.276

Notes: In parentheses are the Newey and West (1987) heteroskedasticity and autocorrelation consistent t-values. Column h is based on estimation for 1981:Q1 to 2001:Q4-h, except for the regressions involving real interest rates, which start in 1982:Q1. The estimates of the intercept term are not shown.

3.2.   The benchmark model

To assess the forecast performance of different models a benchmark is necessary. In finding this model, a univariate autoregressive specification of quarterly real GDP growth is considered for the period 1982:Q1 to 2001:Q4. Starting with a lag length of eight, insignificant coefficients are successively set to zero. This exercise results in the following equation:

$$(9) \qquad \Delta_1 y_t = \underset{(3.31)}{0.0033} + \underset{(1.78)}{0.198} \Delta_1 y_{t-1} + \underset{(1.77)}{0.198} \Delta_1 y_{t-3},$$

t-statistics in parenthesis; adj. R^2: .060, DW: 2.047, S.E: .0049, Portmanteau $\chi^2(16)$: 14.642 (.551), normality-$\chi^2(2)$: .671 (.715), serial correlation $\chi^2(4)$: .489 (.744), ARCH(1): 3.878 (.053), functional form $\chi^2(2)$: .801 (.453), predictive failure ($\chi^2(12)$): .657 (.903).

where Portmanteau $\chi^2(16)$ is the Ljung Box test of autocorrelation for the first 16 lags, normality-$\chi^2(2)$ is the Jarque-Bera test for normality based on skewness and kurtosis of residuals, serial correlation $\chi^2(4)$ is the Lagrange Multiplier test of residual serial correlation for the first four lags, ARCH is the autoregressive conditional heteroskedasticity test, where one lag is considered, functional form $\chi^2(2)$ is Ramsey's RESET test where two terms are taken into account and the test of adequacy of predictions is a Chow-test for the last 12 quarters (1999:Q1 to 2001:4). It is apparent that the diagnostic statistics are not significant at the 5 % test level. However the adj. R^2 is unacceptably small. Therefore, a specification in annual growth rates is considered. Once more, insignificant coefficients are set to zero. The preferred model for the whole sample is:

$$(10) \qquad \Delta_4 y_t = \underset{(2.78)}{0.0036} + \underset{(16.22)}{1.051} \Delta_4 y_{t-1} - \underset{(3.58)}{0.534} \Delta_4 y_{t-4} + \underset{2.70}{0.325} \Delta_4 y_{t-5},$$

t-statistics in parenthesis; adj. R^2: .805, DW: 1.982, S.E: .0055, Portmanteau $\chi^2(16)$: 8.832 (.940), normality-$\chi^2(2)$: 4.870 (.088), serial correlation $\chi^2(4)$: .115 (.977), ARCH(1): 2.005 (.161), functional form $\chi^2(2)$: 2.560 (.084), predictive failure ($\chi^2(12)$): .662 (.781).

The diagnostic statistics are also not significant at the 5 % test level and thus give no hint to any misspecification of equation (10). In comparison to the former equation, the adj. R^2 is considerably higher. Therefore equation (10) is our benchmark model. The intercept in (10) - together with the autoregressive coefficients - is consistent with a trend real GDP growth of 2 – 2.5 % over the period under consideration.

3.3.   Different VARs

We assess the predictive content of M1 relative to other different classes of VARs, including (unrestricted) VARs in terms of the level and of the time difference of the data, cointegrated VARs, or VEC models, and Bayesian VARs (based on the Minnesota prior). For model selection, we do not pre-test the time series properties of individual variables. The preferred model specification is rather selected on the basis of its forecasting performance.[19] According to Sims, Stock, and Watson (1991), the existence of unit roots does not create problems for the estimation of VARs in terms of the level of the variables. Furthermore, due to the low power of univariate unit root tests, an appropriate approximation of the time series properties of the data may be better achieved in the context of multivariate cointegration tests on the respective VAR. If forecasting is the main topic of interest, Clements and Hendry (1998) have shown that VARs in differences may do a good job. Finally, in the context of Bayesian VARs, the Minnesota prior ensures that to the extent that random walk components are important, they are preserved in the model selection.

3.3.1.   VAR and VEC models

In what follows, we use different VARs to assess the predictive content of M1. We start with unrestricted VARs since these are good empirical representations of economic time series as long as enough lags are included (Canova (1995))

$$(11) \qquad X_t = \Gamma + A_1 X_{t-1} + ... + A_o X_{t-o}$$

where X_t is the vector of endogenous variables, Γ the matrix of deterministic terms, especially the intercept term and a linear deterministic trend, A_1 to A_o are the symmetric coefficient matrices and o the selected lag order of the VAR. If the variables are not stationary but cointegrated, the VAR is reparameterised as a vector error correction (VEC) model. The rank of the long run matrix is equal to the number of independent cointegrating relationships. If the variables are not cointegrated, the VEC becomes a VAR in first differences.

The selection of the lag order o is based on the information criteria of Akaike (AIC) and Schwarz (SC) (see Lütkepohl (1993)). Ng and Perron (2001), analyse the AIC and SC and show that it is

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[19] Selection of an empirical model by its forecast performance may be a good way to select a forecasting model, but it is not so to select a model for evaluating economic theory or a policy model (see e.g Clements and Hendry (1998)).

necessary to hold the effective sample size fixed across models to be compared. In the present study, a maximum lag order of 7 is considered and the test period is 1982:Q1 to 2001:Q4. Corresponding to the benchmark model, the values of the criteria are additionally determined for VARs with lags 1 and 4 as well as 1 and 5. We select the lag specification at which a criterion obtains its minimum. For the chosen specification the existence of (no-) autocorrelation is tested by a Lagrange-Multiplier test for autocorrelation of 1 to 8. Moreover, the cointegration hypothesis is checked using Johansen's trace test (Johansen (1995, 2000); Johansen and Juselius (1990)) on the assumption that the intercept term is unrestricted.

3.3.2. A Bayesian VAR

From a Bayesian perspective VARs have been tailored too much towards fitting historical data eventually leading to an overfitting problem. While, in principle, Bayesian VARs (BVARs) comprise as many coefficients as unrestricted VARs, the influence of the data on them is reduced by a statistical procedure to revise prior beliefs in light of the empirical evidence (see, e.g. Todd (1984)).

As the construction of a complete normal prior on a VAR is intractable due to the number of coefficients, the Minnesota prior (Doan, Litterman and Sims (1984)) uses a general prior involving only a few hyperparameters. It has the characteristic feature that the priors on deterministic components are flat, the priors on lags of endogenous variables are independent normal and the means of prior distributions of all coefficients are zero with the exception of the first lag of the dependent variable in each equation which has a prior mean of one. Longer lags have smaller variances around their prior means than shorter lags. Therefore, cross-lag variances have the same relative sizes as the coefficients of own lags. Thus, the prior initially incorporates the assumption that individual time series predominantly evolve like random walks.

3.4.   Results from the VAR models

In the multivariate analysis different three-dimensional and four-dimensional variable sets are investigated. Table 2 below shows the variable sets under consideration.

All three-dimensional processes include the real output and a monetary aggregate (in most cases *m1r*). In addition, an interest rate measure is taken into account. Moreover, the four-dimensional processes contain one of the six other variables like oil prices or exchange rates. The combinations from Table 2 yield 34 different processes which, in combination with the four model classes, produced 135 models. To save space only the results of selected models and processes are presented. The others are available from the authors upon request.

Table 2: Different variable sets for the three- and four-dimensional processes.

Three-dimensional processes contain the variables:		
First	Second	Third
y	m1r	il-is, il, is, il_ger-is_ger, il_ger, is_ger, clr-is, is-i_{m1}, il-i_{m1}, il$_{real(PGDP)}$, il$_{real(HICP)}$, is$_{real(PGDP)}$, is$_{real(HICP)}$
	divr	il-is, il, is

Four-dimensional processes contain the variables:			
First	Second	Third	Fourth
y	m1r	il-is	stock, oilp, oilp€, e, effn, effr
		is	
		il	

The choices of the lag order selection procedures within unrestricted VARs in levels are provided in Table 3. Using AIC, the lags 1 and 4 or 1 and 5 are selected (column 2, no lags in-between). Nevertheless, the LM-test indicates that the null of no autocorrelation is often rejected at the 5 percent test level for the fifth autocorrelation (column 3). The SC in most cases chooses a lag order of 1 (column 4). For this specification, the null of no autocorrelation is rejected in more cases (column 5). In the context of VEC models, the cointegrating properties are tested for the specifications selected by AIC (column 2 with the test results in column 6) and SC (test results in column 8). If both criteria select the same lag order the specification is given in Table 3. In some cases, the null of no cointegration is rejected for the AIC specification. As the test gives no evidence for a cointegrating rank of two, this implies that no stationary process is considered. [The LM-test of no autocorrelation presents evidence that the specification of one cointegration rank reduces the probability to reject the null of no cointegration (column 7 compared to column 3).] For the SC specification the null of no cointegration is mostly rejected (column 8). This specification implies autocorrelation in the estimated residuals (column 9). Therefore, the preferred VEC models are selected by the AIC.[20]

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[20] Mills (1999), p. 36, shows that although theoretically the SC has advantages over the AIC, it would seem that the latter selects the preferred model on more general grounds.

As the results of these lag order selection tests are not unambiguous, we decided to estimate VARs in levels, VARs in time differences and cointegrated VARS, or VECMs. This procedure should be interpreted as a kind of robustness check of the forecasting results. As we are interested in the forecasting performance of M1 for real activity, we always keep *m1r* and *y* in the models considered.

In the context of BVARs, the lag-selection is not as crucial as in unrestricted VARs which build on "hard-shape" exclusion restrictions cutting off the lag length at a certain point. Therefore, the prior builds on an initial lag length which is rather generous (in our case 5 lags with quarterly data). This approach has been maintained for all BVARs presented. The prior is usually tightened with the lag length by choosing a certain decay where we choose a harmonic decay with coefficient 2. This specification gives some improvement over models without lag decays. The BVARs have been set up in terms of levels, including a drift term and with asymmetric priors. These priors reflect the belief that real M1 and the yield spread are more important in affecting GDP than the other way around, and that real M1 is relatively more important in determining GDP than the yield spread. Conversely the remaining relative weights have been set close to zero in the weighting matrix.

Table 3: Lag order estimates of VAR in levels, cointegration tests and rejections of the LM-autocorrelation test

Process with variables	AIC lag order	LM-Test	SC lag order	LM-test	Cointe-gration Rank AIC	LM-test for $r=1$	Cointe-gration rank SC	LM-test for $r \geq 1$
Y, m1r, il-is	1, 5	3	1	1,2	0	1,2	0	1,2,3,4,5
Y,m1r,il	2	-	1	1,5	1	-	1	1
Y,m1r,is	1,5	5	1	1,2,5	1	-	1	1,2
Y,m1r,il_ger-is_ger	1,5	5	1,5	5	0	-	o=1, 1	1,2,3,4,5
Y,m1r,clr-is	1,5	-	1	2	0	2	1	1,2,5
Y,m1r,is-i$_{m1}$	1,5	5	1	1,2	0	-	1	1,2
Y,m1r,il-i$_{m1}$	2	-	1	1,5	1	-	1	1,5
Y,m1r,il$_{real(PGDP)}$	1,4	-	1	1,4,5	0	1	1	1,2,4,5
Y,m1r,il$_{real(HICP)}$	1,4	5	1	1,2,5	0	2,3	1	1,2,3,4,5
Y,m1r,is$_{real(PGDP)}$	1,5	5	1	5	0	5	1	1,3,5
Y,m1r,is$_{real(HICP)}$	1,4	1	1	2,3,5	0	1,3	0	1,3
Y,divr, il-is	1, 5	1	1	1	1	-	1	1,2,3
Y, m1r,il-is,stock	1, 5	1,5	1	1,2,5	1	-	1	1,2,5
Y,m1r,il-is,oilp	1	1,5	1	1,5	o=2, 1	5	1	5
Y,m1r,il-is,oilp€	1	1,5	1	1,5	o=2, 0	5	1	5
Y,m1r,il-is,e	1,5	1,2,3	1	1,5	0	1,2	0	1,2
Y,m1r,il-is,eeffn	1,5	1,2,3	1	1,5	1	2,3	1	1,2
Y,m1r,il-is,eeffr	1,5	2,3	1	1,2,3	0	1,2,8	1	1,2,3

Notes: AIC: Akaike information criterion, LM-test: Rejection of the Lagrange-Multiplier test of autocorrelation for lag order i, where $i = 1,2,…,8$. The cells of the respective column contain the lag orders o, where the null hypothesis of no autocorrelation at this lag is rejected at the 5 percent significance level. SC: Schwarz-criterion, Cointegration rank AIC: Cointegration rank test of Johansen, where the lag order of the corresponding unrestricted VAR is determined by the AIC. If the AIC (SC) lag order estimates are identical to the SC (AIC) lag order estimates, the lag order is augmented by one (is set to unity). Cointegration rank SC: Cointegration rank test of Johansen, where the lag order of the corresponding unrestricted VAR is determined by the SC. Test period is 1982:Q1 to 2001:Q4 and 1983:Q1 to 2001:Q4 for the test regressions involving real interest rates.

The out-of-sample forecasts are computed with a recursive regression method.[21] A recursive estimation of the system yields a series of out-of-sample forecasts for the different forecasting horizons $h=1,\dots,8$. The starting coefficients are computed over the period 1980:Q1 to 1993:Q4. Using these coefficients, the first forecasts are determined. The forecast errors are the differences between the forecast of y and the historical values of y. In a next step, the sample is extended by one quarter and the system is re-estimated to calculate the forecasts again. This procedure is continued until the end of the sample. It is conducted for the benchmark model (10) and the different VAR models.

The accuracy of forecasts can be judged by various statistics about the forecast errors. In this study the root mean square forecast errors (RMSFE) are used. To assess the relative predictive accuracy of two forecasting models, the Diebold Mariano test is selected, which has an asymptotic normal distribution (see Diebold and Mariano (1995)).[22] The longest interval for all forecasts is from 1994:Q1 to 2001:Q4, hence the maximum length of the forecast period is 32. The values of the RMSFE for real output of the benchmark model are given in the first row of Table 4. In general, they increase with the forecasting horizon. The results of the other approaches are all given relative to this benchmark (RMSFE of the alternative model divided by RMSFE of the benchmark model). Values greater than one indicate that the alternative model is worse than the benchmark model.

Considering VARs in levels with or without a deterministic trend all models are worse than the benchmark (not shown, but available upon request). Very often the differences are even significant at the 5 or 1 percent level. Up to forecast horizon 3 the second-best model after the benchmark model is the one with the additional variable (*clr-is*), for the horizons 4 to 6 the model with *is* and *oilp* and for the two longest horizons 7 and 8 the model with the Divisia aggregate and *il*. It is worth noting that the differences are slightly smaller if VARs in levels without a deterministic trend are estimated. However, in any case this alternative does not dominate the benchmark either. The results of the cointegrated VARs, or VEC models, point in the same direction (not shown, but available upon request). Often the results are better than the results of the VARs in levels. Nevertheless, only for $h=1$ some variants of this model class outperform the benchmark. But the difference is in no case statistically significant at conventional significance levels.

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[21] A comparison of out-of-sample and in-sample tests of predictability may be found in Inoue and Kilian (2002).

[22] See the appendix for a description of the test.

Table 4: Root mean square forecast errors for VARs in first differences in relation to the root mean square forecast error of the benchmark equation

Model/Process with variables	Spe-cifi-cation	Forecast horizon h							
	O	1	2	3	4	5	6	7	8
Benchmark	1,4,5	.0040	.0062	.0082	.0097	.0112	.0126	.0124	.0124
y, m1r, il-is	1	0.829#	0.837+	0.790*	0.781*	0.734*	0.750*	0.841+	0.844#
y,m1r,il	1	0.787+	0.796+	0.731*	0.751*	0.720*	0.743*	0.848+	0.836#
y,m1r,is	1	0.813#	0.826+	0.776*	0.778*	0.733*	0.752*	0.846+	0.841#
y,m1r,il_ger-is_ger	1	0.823#	0.829+	0.782*	0.776*	0.734*	0.753 *	0.848+	0.853
y,m1r,clr-is	1	0.815#	0.838#	0.783*	0.784*	0.729*	0.746*	0.833+	0.824#
y,m1r,is-i_{m1}	1	0.807+	0.821+	0.772*	0.777*	0.732*	0.751*	0.845+	0.839#
y,m1r,il-i_{m1}	1	**0.778+**	**0.780+**	**0.704***	**0.732***	**0.709***	**0.738***	0.852#	0.841#
y,divr, il-is	1	0.836#	0.843#	0.820+	0.837*	0.796*	0.798*	0.864+	0.876
y,m1r,il$_{real(PGDP)}$	1	0.815+	0.832+	0.806+	0.808+	0.761*	0.760*	0.815*	0.792*
y,m1,il$_{real(HICP)}$	1	0.831#	0.832+	0.795*	0.790*	0.743*	0.748*	**0.810***	**0.786+**
y,m1,is$_{real(PGDP)}$	1	0.813+	0.823+	0.796*	0.804+	0.762*	0.766*	0.822+	0.803+
y,m1r,is$_{real(HICP)}$	1	0.831#	0.831+	0.791*	0.787*	0.743*	0.750*	0.813*	0.792+
y, m1r,il-is,stock	1	0.828#	0.854#	0.828+	0.818*	0.775*	0.780*	0.849+	0.809*
y,m1r,il-is,oilp	1	0.809+	0.821+	0.783*	0.786*	0.739*	0.750*	0.825+	0.804+
y,m1r,il-is,oilp€	1	0.844#	0.868	0.820+	0.806+	0.748*	0.753*	0.820+	0.796+
y,m1r,il-is,e	1	0.889	0.903	0.871#	0.828#	0.754+	0.757+	0.838+	0.836#
y,m1r,il-is,eeffn	1	0.875	0.911	0.888	0.851#	0.775+	0.766+	0.829+	0.805+
y,m1r,il-is,eeffr	1	0.871	0.904	0.878#	0.845#	0.773+	0.766*	0.831+	0.806+

Notes: The column "Specification" includes the lag order (o) of the VAR in first differences. The cells contain the root mean square forecast error of the process divided by the root mean square forecast error of the benchmark model. #,+,* denote significance at the 10, 5, 1 percent level, respectively. Normal distributed statistic using the critical values 1.64486; 1.95997; 2.55758. The bold values gives the lowest RMSFE for this forecast horizon for all approaches.

The results obtained on the basis of the BVAR models also show that it was not possible to outperform the univariate benchmark model using a BVAR (not presented, but available upon request). This result is in line with recent findings in Canova (2002). Except for one-step-ahead forecasts, BVARs are able to outperform the respective non-Bayesian VARs in levels over all horizons. The best performing model is the one containing the Divisia aggregate. Second to these models are BVARs containing real interest rate measures in addition to M1.

The situation is completely different with respect to difference VARs. Table 4 shows the results of selected VAR models with only variables in first difference (and no long-run relations), where a lag order of one is selected.[23] It is apparent that all models are better than the benchmark model for the whole range of forecast horizons. The reduction in the RMSFE is 20 to nearly 30 percent. The best model includes the variables y, $m1r$ and (il-i_{m1}) for the horizon $h = 1$ to 6 (see the bold figures in Table 4). For $h = 7$ and 8, the best model is y, $m1r$ and $il_{real(HICP)}$. Having in mind that the differences between the benchmark model and the alternative model are not significant if they are less than 10 percent this implies that the choice of the variables has only a small effect. It seems that it is more important to choose the "right" general class of VAR models.[24]

4. Summary and Conclusions

The purpose of this paper was to assess the forecast performance of narrow M1 for real activity, measured by the growth rate of real GDP, in the euro area. After a brief review of the empirical literature, we recall a number of theoretical arguments why money might be useful for real developments beyond the effects of monetary policy captured by a short-term interest rate.

Using a single equation methodology recently proposed by Hamilton and Kim (2002), we find that M1 has important indicator properties with respect to real output, even after controlling for other variables. In contrast to findings for the U.S., the evidence in the euro area seems to suggest that - over the sample period under investigation (i.e. 1981-2001) - M1 outperforms the yield spread in terms of its predictive content for cyclical movements in GDP. These properties are confirmed also when looking at a broader information set comprising non-monetary indicator

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[23] In most cases, where the AIC chooses a lag order of 1 and 4 or 1 and 5 for the level approach, a lag specification of 1 and 4 or 1 and 5 for the VARs in first differences does not yield better results than those presented.

[24] A single equation approach with $\Delta_4 y$, $\Delta_4 y_{t-1}$, $\Delta_4 m1r$ and one or more of the variables taken into account in the Tables 3 to 6 would not be able to outperform the benchmark model with dynamic forecasts irrespective of the forecast horizon considered.

variables. Hence, the findings appear to be quite robust when extending the information set to other relevant variables.

Comparing the out-of-sample forecasting performance of different classes of VAR models with a univariate benchmark model 1 to 8 quarters out, only VARs in first differences are able to outperform the univariate benchmark model and yield significantly better forecasting results at all forecast horizons. This may be due to the fact that the evidence for the existence of cointegration relationships is not unambiguous and that most level variables seem to contain a unit root. As we are only interested in short-term forecasting, economic theory would also suggest that the dynamic part is the essential one. Moreover, as Clements and Hendry (1998, ch. 6 and 7) show, vector autoregressions in first differences are better able to capture structural breaks during the forecast period than VECMs or VARs in levels, even if they are mis-specified. In their view, the use of growth-rate data is one possibility to make forecasts more robust to shifts (and measurement errors) in the levels of variables. Sometimes it is simply the case that one-off innovations affect the levels, but do not persist, so that rates of growth are not affected.

The best models for the shorter horizons (up to six quarters) is the one which - besides real M1 and real GDP growth – includes the spread between the yield on government bonds and the own rate of return of M1. For the two longest horizons (7-8 quarters), the best model additionally takes into account the real long-term interest rate.

One interesting field of future research would be to elaborate in more detail on the theoretical justifications on the indicator properties of M1 for real activity. Empirically, one could analyse which of the components of M1, i.e. currency and overnight deposits, are responsible for the results. Moreover, it may also be interesting to examine the model performance at forecast horizons longer than 2 years or use other metrics as the root mean squared error, e.g. the direction of change or the general forecast-error second moment matrix (Clements and Hendry (1998) ch. 3.6). And finally, as the primary goal of the monetary policy of the ECB is to maintain price stability, it is of overriding importance to assess how changes in demand conditions indicated by movements in M1 affect price developments.[25]

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[25] Leading indicator properties of M1 for prices have also been analysed in Nicoletti Altimari (2001).

Appendix 1: List of Variables and Symbols

ß   discount factor

c   real private consumption expenditure

clr   composite lending rate

divr   real divisia aggregate

$\Delta_{1(4)}$ logarithmic first (fourth) difference

e   US-$/€ exchange rate

eeffn   nominal effective exchange rate of the €

eeffr   real effective exchange rate of the €

E   expectations operator

ε   elasticity

h   forecast horizon

HICP   Harmonised Index of Consumer Prices

i   interest rate

is   short-term interest rate

il   long-term interest rate

i_m own rate of interest of money

i_{ger} interest rate in Germany

i_{real} real interest rate

m   real money balances

m1r   real M1 calculated with the GDP deflator

oilp   oil prices (in $)

oilp€   oil prices (in €)

pgdp   GDP deflator

stock   world stock prices

u   per-period utility

y   real income, real GDP

Appendix 2: The Diebold-Mariano Test

The accuracy of forecasts can be assessed by various statistics about the forecast errors. In this study, the root mean square forecast errors are selected. To check the relative predictive accurarcy of two forecasting models, different test statatistics are suggested and analysed by Diebold and Mariano (1995). Their preferred test statistic is

$$\hat{d}_F = F^{-1/2} \frac{\sum_{t=T+1}^{S-h}(\hat{e}_{0,t+h}^2 - \hat{e}_{1,t+h}^2)}{\hat{\sigma}_F}$$

where T denotes the length of the estimation period, F is the length of the forecast period, hence $S=T+F$, $h \geq 1$ is the forecast horizon, $\hat{e}_{0,t+h}^2$ and $\hat{e}_{1,t+h}^2$ are the squared forecast errors of the benchmark model and the alternative model, respectively, using consistent estimators, and

$$\hat{\sigma}_F = \frac{1}{F} \sum_{t=T+1}^{S-h}(\hat{e}_{0,t+h}^2 - \hat{e}_{1,t+h}^2)^2 + \frac{2}{F} \sum_{j=1}^{l_F} \omega_j \sum_{t=T+1+j}^{S-h}(\hat{e}_{0,t+h}^2 - \hat{e}_{1,t+h}^2)(\hat{e}_{0,t+h+j}^2 - \hat{e}_{1,t+h+j}^2),$$

where $\omega_j = 1 - \dfrac{j}{l_F + 1}$, $l_F = o(F^{1/4})$. The parameter ω_j is the Bartlett weight and l_F is the truncation parameter depending on the converging rate of F. For $F = 32$ we choose $l_F = 2$. The test statistic is denoted the Diebold-Mariano (dm) test. The null of equal predictive ability is

$$H_0 : E(e_{0,t+h}^2 - e_{1,t+h}^2) = 0$$

while the alternative is

$$H_1 : E(e_{0,t+h}^2 - e_{1,t+h}^2) \neq 0$$

Under the null hypothesis, this statistic has an asymptotic standard normal distribution. Harvey, Leybourne and Newbold (1997, 1998) analyse the test statistic using an extensive Monte Carlo design, and find that the test has good size and fairly good power properties.

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