BANCA D'ITALIA

Temi di discussione

del Servizio Studi

Money Demand in the Euro Area: Do National Differences Matter?

by L. Dedola, E. Gaiotti and L. Silipo

Number 405 - June 2001

MONEY DEMAND IN THE EURO AREA: DO NATIONAL DIFFERENCES MATTER?

by Luca Dedola[*], Eugenio Gaiotti[*] and Luca Silipo[*]

Abstract

This paper assesses the relevance of national information in estimating the demand for euro-area M3 from three perspectives. First, we check whether aggregating national money demands is appropriate. Second, we compare time-series and panel methods to estimate aggregate long-run coefficients. Finally, we investigate the differences among national money demands. We find that the hypothesis of perfect aggregation is not rejected. Nevertheless, some estimates of area-wide long-run parameters are sensitive to the method used to combine national information. We also find that the main difference among individual countries' money demands is their interest elasticity, as well as the existence of country-specific structural breaks. We conclude that the area-wide equation is an appropriate analytical tool; national information may be useful to interpret "special factors" and institutional events.

JEL Classification: E41, C22, C23.

Keywords: money demand, aggregation, European Central Bank.

Contents

[*] Bank of Italy, Research Department.

1. Introduction[1]

The strategy of the European Central Bank assigns a "prominent role" to money, signalled by the announcement of a reference value for M3 growth.[2] This has led the properties of money demand in the euro area to become the focus of new interest and the subject of a number of empirical studies.

As the policy debate shows, having a reliable estimate of money demand is important. The stability of money demand determines whether money is an appropriate guide to policy. The magnitude of its income elasticity determines whether there is a trend in the velocity of circulation, which is relevant in determining the reference value for money growth. The interest rate elasticity of money demand, in turn, is essential to interpreting the short-run movements of money around its reference value.[3] It has also been argued that attention should also be paid to the level of money balances;[4] the practical relevance of the argument depends on the possibility of estimating the equilibrium level of money balances with reasonable accuracy.

So far, the issue of the properties of money demand has been mostly addressed from an area-wide perspective, using aggregate data at the euro-area level. The information contained in the national "contributions" to M3 has been largely neglected; only occasional reference has been made to "special factors" (of a structural or institutional nature) affecting the national contributions to M3. This approach contrasts with the treatment of other macroeconomic variables. For instance, area-wide price and output projections are obtained making large use of both national and area-wide econometric models.[5]

---

[1] We thank D. Focarelli, F. Busetti, R. Golinelli and the participants in seminars at the Bank of Italy and at the Graduate Institute for International Studies in Geneva for their useful comments and suggestions. We are grateful to Claudio Trevisan for his skill and dedication in the construction of the statistical database. The usual disclaimer applies.

[2] See European Central Bank [1999a].

[3] "A deviation of monetary growth from the reference value will prompt further analysis to identify and interpret the economic disturbance that caused the deviation" (European Central Bank [1999a]).

[4] E. g., European Central Bank [1999b]; Trecroci and Vega [2000].

[5] The Eurosystem's projections are conducted by ECB and national central banks experts; they are "obtained in a way that is consistent with individual country assessments, incorporating the full range of expertise available". See European Central Bank [1999a].

Our aim is to fill this gap. Building on previous work by a number of authors, we seek to offer a contribution in two respects. Firstly, we compare different methods of estimating the parameters of area-wide money demand starting from individual countries, since estimates of "average" coefficients obtained from aggregate time series may give inconsistent and potentially misleading results. Secondly, we discuss differences in the properties of "demand" functions for national contributions to M3; national information may be relevant for policy when there are cross-country differences in behavioral equations.

From a methodological standpoint, a feature of our approach is the use of a newly constructed series of the opportunity cost of holding money. It has been shown for a number of European countries that estimates of interest rate elasticity are sensitive to the introduction of the own rate. Yet, the literature on money demand in the euro area so far has not computed or used own rates of return on monetary aggregates, and this omission has resulted in the estimation of a barely significant reaction of money demand to interest rates.

The paper is organized as follows. Section 2 summarizes the debate on money demand in the euro area. Section 3 discusses in more detail the issues we intend to tackle. Section 3 presents our dataset. In Section 4, we estimate a benchmark aggregate money demand equation for the euro area, including the own rate among the regressors and comparing the results with previous research. Section 5 assesses the appropriateness of aggregation on the basis of tests or criteria existing in the literature; section 6 compares different methods of computing aggregate coefficients from individual countries' data; section 7 discusses the individual countries' results and their implications.

2. Money demand in the euro area: what do we know?

Research on the demand for money in the euro area is relatively recent, but the literature is already very large. The most recent contributions, by Coenen and Vega [1999], Golinelli and Pastorello [2000], Brand and Cassola [2000], make use of the official M3 series as defined by the Eurosystem in 1998. Previous work is mostly based on the pre-EMU definition of "harmonised M3" (M3H), which was agreed upon for comparisons by the central banks of the EU prior to EMU (a definition similar, but not identical, to current

M3).[6] All previous research points to the existence of a stable area-wide money demand function linking real money balances to output and interest rates, a result that forms the basis of the important role attributed to M3 in the ECB's strategy.

Most results indicate that the income elasticity of money demand is higher than one.[7] Income elasticity higher than one implies declining velocity; it determines what is the appropriate long-run rate of money growth, i. e. the rate consistent with long-run price stability and potential output growth. The ECB's reference value for money growth (currently 4.5 percent on an annual basis) is estimated as the sum of the inflation rate implicit in the definition of price stability,[8] potential output growth[9] and the trend change in velocity, which, on the basis of the above-mentioned results, is assumed to be an annual decline of 0.5/1.0 percent.[10]

As far as the interest rate variable is concerned, both long and short-term rates are usually included in the European demand for money. In principle, risk-free short rates should enter the demand for transaction money (Ando and Shell [1975]); long rates may be relevant, however, in presence of market imperfections and when money also serves as a store of wealth (Baba, Hendry and Starr [1992]).

A peculiar result of many recent estimates is the small response of euro area money demand to a hike in interest rates. According to Coenen and Vega [1999], money demand is inversely related to the slope of the yield curve (the difference between long- and short-term rates); they present the result by interpreting the short-term rate as a proxy for the own rate of return on money. However, the result has implications for policy experiments: in ordinary

---

[6] See Angeloni, Cottarelli and Levy [1992], Monticelli and Papi [1996], Fagan and Henry [1998], Fase and Winder[1998].

[7] The conclusion is reached by Coenen and Vega [1999], Fagan and Henry [1998], Golinelli and Pastorello [2000], Brand and Cassola [2000] for the euro area; a survey of results for individual countries is in Filosa [1995]. By contrast, Fase and Winder[1998] include financial wealth in the European demand for money and find that income elasticity is less than one.

[8] The ECB defines price stability as "a year-on-year increase in the Harmonised Index of Consumer Prices (HICP) for the euro area of below 2 per cent", to be maintained over the medium term.

[9] In the latest review of the reference value, potential growth in the euro area was assumed to be between 2 and 2.5 percent (European Central Bank [2000b]).

[10] See European Central Bank [1999b] and [2000b].

circumstances it would imply a perverse response of money (i. e., an increase) to a monetary policy tightening, as the yield curve usually flattens after an increase in policy rates. Similarly, Fagan and Henry [1998] obtain a fairly low interest rate elasticity, with a positive sign for the short-term rate and a negative one for the long-term rate. They also argue that the short-term rate could be "picking up the effect of the 'own' rate on money while the long-term rate is acting as a measure of the opportunity cost". Given the high collinearity of interest rates, such a misspecification is not likely to change the fit of the equation, but it has important implications both for the "controllability" of money[11] and for extracting information from its short-run movements.[12]

Omission of the own rate may result in a misspecified equation, which could produce incorrect answers to some policy questions. The estimates of Filosa [1995] show that, for all major European countries, the omission of the own rate from the equation does result in estimated coefficients being positive for short-term rates and negative for long-term rates. However, when a proper measure of the own interest rate is included as a separate explanatory variable, the estimated coefficients of the own rate on money are positive and statistically significant while the returns on alternative assets (both short and long) have the correct negative coefficients.

Recent papers, such as Golinelli and Pastorello [2000] and Brand and Cassola [2000], use the long rate, not the slope of the yield curve, as a measure of the opportunity cost, obtaining a somewhat larger elasticity. The latter authors argue that the short rate is not an appropriate proxy for the own rate, while "there is a strong resemblance between the dynamics of the long rate and the spread constructed using the own rate". In the following, we introduce for the first time an own rate measure for the euro area as well as for the individual countries. The issue of the construction and introduction of the own rate in a money demand function for the euro area has also been recently addressed by Calza *et al.* [2001].

[11] However, controllability of M3 is not considered an important requisite for the ECB's strategy. See European Central Bank [1999a].

[12] According to the theoretical literature (Friedman [1990]), movements in money due to a response to interest rate changes have no information value for income or inflation and should be ignored in setting the policy instruments.

3. E pluribus unum? A reappraisal of the aggregation issue

In what follows, we focus on testing aggregation, assessing the robustness of estimates of structural coefficients, and checking the differences in national money demand equations.

Testing aggregation. An OECD report (OECD [2000]) extensively discusses the pervasive aggregation problems in estimating euro-area-wide behavioral equations, with a particular focus on wage equations. It concludes that aggregation should be used carefully, as cross-country variation does matter, area-wide influences on country-specific variables are likely to be small, and institutional and policy differences are much larger in the European Union than in federations, raising the aggregation bias. However, the report maintains that these problems do not apply to money demand.

The issue of whether aggregation of national equations is legitimate was addressed in the early literature on money demand in the euro area focusing on the relative importance of *aggregation bias* (the bias introduced by aggregating individual equations when aggregation conditions are not met) versus *specification bias* (national money demand functions may be misspecified due to the omission of area wide variables). The latter case may apply if the demand for national money also depends on foreign variables, as would be the case if there were currency substitution within European portfolios. The issue was addressed comparing the standard errors of the aggregate equation with those of each national equation. Since the standard error of the aggregate equation always turns out to be smaller than that of individual equations, the common conclusion is that aggregation bias is not likely to be a major problem (Monticelli and Papi [1996]; Fagan and Henry [1998]).[13]

However, this conclusion is not warranted. The fact that a macro equation has a smaller standard deviation (a higher R^2) than a micro equation is not relevant in judging the performance of either equation (Grunfeld and Griliches [1960]). The relevant comparison is

[13] Based on cross correlations of the individual countries' equations, Fagan and Henry [1998] also suggest that currency substitution is not the major factor behind the superior performance of the euro-area equation; rather, the latter is mostly due to statistical averaging of individual disturbances. Arnold [1994] had argued that the stability of money demand in the euro area is a statistical artifact and that it is bound to disappear as soon as domestic shocks became positively correlated as a result of financial convergence. The computations of Fagan and Henry [1998], based on the estimated covariance matrix of the national disturbance, show that these effects are small however, even if high positive correlations should arise owing to EMU.

whether the aggregate equation explains aggregate data better than all the national equation combined. In this respect, it is interesting to apply the existing procedures to check for aggregation to European money demand. The most frequently used (as the Grunfeld and Griliches [1960] model-selection criterion, the Pesaran, Pierse and Kumar [1989] test of perfect aggregation) have not been employed on euro-area M3.

Assessing the robustness of estimates of structural coefficients. Moreover, research has not considered the effects that aggregation itself may have on the estimates of the structural parameters of money demand, which are relevant for conducting policy experiments. As Pesaran, Pierse and Kumar [1989] and Pesaran, Shin and Smith [1999] make clear, aggregation tests are based on the forecasting performance of the area-wide equation, which does not necessarily coincide with consistency of the estimates of the structural parameters. This may be a problem when national coefficients differ.

Parameter equality is only a sufficient, not a necessary, condition for aggregation. [14] However, Pesaran and Smith [1995] show that in dynamic models where individual coefficients differ not all methods for estimating average coefficients are consistent and some may be seriously misleading. They compare four ways to estimate average coefficients: using aggregate time series, estimating national equations and then averaging the coefficients (the "mean group estimator"), pooling, and running cross-section estimates with long-period averages for each country's variables.

In dynamic models, the mean group estimator gives consistent estimates of the average area parameters. By contrast, the estimates obtained from aggregate time series are inconsistent, unless the coefficients are the same across groups, since in a dynamic model the aggregate disturbance turns out to be correlated with the aggregate regressors in a very complicated way; [15] the same holds for pooled estimators. Pesaran and Smith's solution is to

---

[14] Even if parameter equality is not met, aggregation may be valid owing to the invariance of the composition of the regressors across the individual equations over time; there can be an aggregate demand for money, whose coefficient are the (weighted) average of the domestic coefficients.

[15] Pesaran and Smith [1995] argue that the autocorrelation may be so complex that standard procedures for dealing with it cannot be used. They show that, in general, there is a difference between the "structural" aggregate money demand and the optimal predictor for money (the latter has to take into account a very complicated pattern of serial correlation).

estimate individual micro-relations separately and then explicitly calculate the averages of the estimated parameters.

Checking the differences in national money demand equations. Knowing the differences in national money demand is also of interest *per se*. De Grauwe [2000] and Angelini *et al.* [2000] have shown that national information may be relevant for area-wide policy when there are substantial cross-country differences in behavior. In this case, it may be optimal for the area-wide policymaker not only to look at the area-wide figures, but also to consider national contributions. Extending the intuition from the classical Poole model, the importance assigned by the policymaker to each country's contribution to M3 should increase with the variance of the real disturbances and income elasticity, and decrease with the variance of the money demand disturbances in that country (see appendix I).[16]

The assumption of equality of coefficients of individual countries' money demands was tested and rejected by the early work of Angeloni, Cottarelli and Levy [1992]. More recently, Golinelli and Pastorello [2000], pooling domestic money demand for the Euro-10 countries, impose and test the restriction that the long-run parameters are the same for all countries. Even this weaker restriction is rejected for the whole set of countries. Panel data commonly reject the assumption that coefficients are equal, as Pesaran and Smith [1995] emphasize.

Many past accounts suggest that money demand in the five largest countries in the area differ. An example is given by the contrasting trends in velocity, shown in figure 4: a clearly rising trend in the M3/GDP ratio in Germany and some other countries, and a declining one in Italy and France.

One instance of institutional and structural differences in Europe is the composition of private portfolios. The alternative asset available to money holders, whose yield enters the money demand equation, depends on the menu of existing financial instruments and on the institutional framework in the estimation period. According to the existing literature,[17] long-

---

[16] Poole [1970] has shown that, to make optimal use of the information contained in money, the response of the interest rate to money should depend on the relative variances of real and monetary disturbances and on the interest elasticity of money demand.

[17] For a survey, see again Filosa [1995]; for Spain, see Vega [1998].

term rates represent the yield on the alternative asset in Germany and the Netherlands, while short-term rates are more relevant in Spain and Italy. Also, country-specific institutional events affected money demand in the past two decades. The observed episodes of instability are largely uncorrelated across countries. An instance is the 1989 reunification in Germany. In Italy and Spain, structural shifts out of money occurred in the first half of the 1980s and after 1992, owing to various changes in the supply of alternative financial instruments (in Italy, a shift to T-bills in the early 1980s and a shift to mutual funds at the end of the 1990s). In the Netherlands, here was a sharp increase in the demand for money of corporations and financial institutions towards the end of the 1980s. By contrast, the long-run income elasticity is usually estimated to be more similar across countries; nonetheless, it is usually found to be larger than one in Spain and Germany, smaller than one in Italy and about one in the Netherlands.

4. The data and a benchmark aggregate money demand

Our dataset includes real money (defined as the log-difference between M3 and the consumer price index), GDP, long and short interest rates, consumer price inflation and the own rate of return on money. All data refer both to the euro area and to the member countries (figures 1- 4)[18]. A full description of the data is given in appendix II. A few remarks are necessary here.

The own rate of return of money was reconstructed based on the available evidence for each country on the yield on three categories of instruments included in M3: currency, deposits and marketable instruments. When necessary, separate data on yields for "overnight deposits", "deposits with agreed maturity up to 2 years" and "deposits redeemable at notice up to 3 months" were also used. After 1990, most data come from the data falling into these categories; before that date, other national sources were used.

By construction, the national contributions to M3 sum exactly to area-wide M3, by construction. However, money demand equations are specified in log-linear form; the linear

[18] Greece entered EMU in January, 2001. Due to data availability constraints, our estimates only include the eleven countries that entered in January 1999.

aggregation of the log of the M3 components is not equal to the log of area-wide M3. For this reason, Golinelli and Pastorello [2000] choose not to conduct formal aggregation tests on individual countries' money demands, as most of the latter are based on the assumption of linear aggregation. We exploit the fact that there is little difference between the behavior of the log of area-wide M3 and of a weighted average, with proper weights, of the log of national components (see Fagan and Henry [1998]. That is,

$$\log(M_t) \approx \sum_{k=1}^{11} w_k \log(M_{k,t}) + c$$

(where the subscript k stands for the country and w_k is its constant GDP share). Figure 5 compares the two series. Formal aggregation tests can be conducted on $\Sigma w_k \log(M_{k,t})$ instead than $\log(M_t)$.

Unit root tests for real money balances, real GDP, the (month-to-month) inflation rate and the differentials of both long and short-term rates vis-a-vis the own rate on money are presented in table 1. The ADF statistics show that almost all the variables are I(1) (only in a few cases is the assumption that interest differentials or the inflation rate are I(1) rejected).

According to pairwise cointegration tests (not reported), for most countries stationary combinations exists for each pair of two rates; the own rate enters the cointegrating equations with coefficients different from 1, consistently with the fact that the long-own rate spreads and the short-own rate spreads are I(1).

We estimated an aggregate money demand for the euro area, primarily to have a benchmark against which to evaluate the individual countries' data in section 5. Given our objective, we mostly work in a single equation framework, referring to previous work (Brand and Cassola [2000], Coenen and Vega [1999], Golinelli and Pastorello [2000]) for the derivation of a money demand equation from a system approach.

The identification of structural long-run money demand is not straightforward, since the cointegrating relations among the variables are potentially more than one. Golinelli and Pastorello [2000] conclude that a structural relation linking money, income and either the long or the short-term interest rate holds, with coefficients of the expected sign; they find that inflation does not enter that relation. Key elements of their conclusion are that i) excluding the inflation rate from the long-run vector does not disrupt cointegration and ii) inflation enters a simpler (structural) cointegrating relation with the long run rate.[19] They interpret these results as indicating that, even if a cointegration vector including money, income, interest rates and inflation is found, this does not need to be the structural long run demand for money. From a theoretical standpoint, money demand should not depend on inflation directly, but indirectly through its effect on nominal interest rates. A similar result is reached by Brand and Cassola [2000] in a demand system including real money, inflation, output, long and short rates. They identify three long-run relationships: the Fisher equation, the long-short spread and money demand.

Our main departure from these papers is the assumption that the differential between market rates and the own rate on money is the measure of the opportunity cost that enters the long-run relationship. We introduce this hypothesis a priori, based on economic theory. Using the two spreads (long-own rate, short-own rate) gives appreciable results, with the expected (negative) sign. Table 2 reports the results of a series of Johanssen cointegration tests, showing that a cointegrating relation exists, and it includes the long-term differential. Simple static regressions also confirmed that inflation is not needed to find a stationary combination of the variables.

We then estimated a dynamic, single-equation model of money demand. Our benchmark specification includes four lags of real money, output, the two differentials and the change in inflation (plus seasonal dummies). We include both interest rate differentials mostly for a comparison with the disaggregate estimates in the following section, as the role of short versus long rates can be an important difference across countries.

A version of this model, after deleting most statistically not significant terms following a general-to-specific procedure, is shown in table 3. The long-run interest rate differential enters the equilibrium relation in a statistically significant way.[20] Diagnostics tests are satisfactory. Stability tests are also satisfactory (Figure 6), although some sign of instability

[19] They use the concept of "irreducible cointegrating relationship" suggested by Davidson [1998], to find the structural long run relationship for money demand for the euro area.

[20] We retained the short-term differential in the equilibrium condition, even thought it was non-significantly different from zero, for the purpose of comparison with the results in the following section.

are present at the beginning and at the very end of the sample; this leaves the issue open of whether the start of EMU has permanently changed the properties of money demand. Table 4 compares the results with those of Golinelli and Pastorello [2000], Coenen and Vega [1999], Brand and Cassola [2000]. A comparison is also reported with the cointegrating vector we obtained from the Johanssen procedure mentioned above. Overall, the inclusion of the interest rate differential seems to increase the interest elasticity of money demand. [21]

5. Should we aggregate?

Before discussing the estimates of the structural coefficients, we assess the appropriateness of aggregation from the standpoint of the overall fit of the equation. To this end, we run two sets of estimates: the first is the aggregate equation discussed in the previous section:

$$(1)\quad \begin{aligned} \Delta\log(\frac{M_t}{P_t}) &= a+\sum_{i=1}^{3} b_i \Delta\log(\frac{M_{t-i}}{P_{t-1}}) + \sum_{i=0}^{3} c_i \Delta\log Y_{t-i} + \sum_{i=0}^{3} d_i \Delta(rl-rm)_{t-i} + \sum_{i=0}^{3} f_i \Delta(rs-rm)_{t-i} + \sum_{i=0}^{3} g_i \Delta p_{t-i} \\ &+ j\,[\log(\frac{M_{t-1}}{P_{t-1}}) - a\log Y_{t-1} - b\,(rl-rm)_{t-1} - g\,(rs-rm)_{t-1}] + e_{a,t} \end{aligned}$$

(where M/P is real money, Y is GDP, rl, rs, rm are the long, short and own rate on money).

A second set of regressions is obtained by estimating (1) separately for each country k:

$$(2)\quad \begin{aligned} \Delta\log(\frac{M_{k,t}}{P_{k,t}}) &= a_k+\sum_{i=1}^{3} b_{k,i} \Delta\log(\frac{M_{k,t-i}}{P_{k,t-i}}) + \sum_{i=0}^{3} c_{k,i} \Delta\log Y_{k,t-i} + \sum_{i=0}^{3} d_{k,i} \Delta(rl-rm)_{k,t-i} + \sum_{i=0}^{3} f_{k,i} \Delta(rs-rm)_{k,t-i} \sum_{i=0}^{3} g_{k,i} \Delta p_{k,t-i} \\ &+ j_k[\log(\frac{M_{k,t-1}}{P_{k,t-1}}) - a_k\log Y_{k,t-1} - b_k(rl-rm)_{k,t-1} - g_k(rs-rm)_{k,t-1}] + e_{k,t} \end{aligned}$$

We then apply the Grunfeld and Griliches [1960] prediction criterion for aggregation, according to which the disaggregate model should be chosen if

[21] We also ran a number of static long-run regressions for money demand, including long, short and own interest rates. The inclusion of the own rate of return appears useful in determining the “correct” sign for both short and long rates. Without this variable, we find a positive sign on the short rate and a negative sign on the long-run rate. In this case, one would – mistakenly – conclude that the slope of the yield curve is the “correct” variable that enters money demand; when the own rate is included, the signs are usually what one would expect. This is in line with the findings of Filosa [1995].

(3) $$s_d^2 < s_a^2$$

Here, $\sigma^2_a = var(e_{a,t})$, $\sigma^2_d = var(e_{d,t})$ and

(4) $$e_{d,t} = \sum_{k=1}^{m} w_k e_{k,t}$$

where w_i is the i-th country GDP weight. Asymptotically, (3) will always be satisfied unless the individual equations are mis-specified (in the money demand case, this could be a consequence of currency substitution affecting national money demands, but canceling out in the aggregate).

Table 5 shows that the standard error of the estimate based on the disaggregate model (σ^2_d) is smaller - but only slightly - than the standard error of the aggregate equation (σ^2_a),[22] a conclusion somewhat in contrast with previous results in the literature. According to the Grunfeld-Griliches criterion, the disaggregate model can be chosen. However, the difference is very small, suggesting that the worsening of the fit deriving from aggregation is negligible. To properly assess it, a formal testing procedure is needed.

The Grunfeld-Griliches criterion is not a formal test for the conditions for valid aggregation. In order to provide a more stringent analysis of the feasibility of aggregation for our model, we flank this simple criterion with a more technical measure of the error made during the aggregation process.

We use the test of perfect aggregation developed by Pesaran, Pierse and Kumar [1989]. The test is based on the comparison between the residuals generated by the two models. The a_m statistics is defined as follows:

(5) $$a_m = m^{-1}(\bar{e}_a - \bar{e}_d)'\overline{\Psi}_m^{-1}(\bar{e}_a - \bar{e}_d) \sim c_n^2$$

where m is the number of cross-sectional units, $\bar{e}_a$ and $\bar{e}_d$ are the (tx1) vectors of residuals generated respectively by the aggregate equation and the disaggregated estimated system

[22] The reduction of the SEE from individual equations (about 0.9 percentage points) to area-wide equations (about 0.3 percentage points) is consistent with the simple effect of averaging eleven residuals with very low cross-correlations.

(the latter defined as in (4) above), and the (txt) matrix $\overline{\Psi}$ is defined as

(6)
$$\overline{\Psi}_m = m^{-1}\sum_{i,j=1}^{m} s_{ij}(\overline{A}_i - \overline{A}_a)(\overline{A}_j - \overline{A}_a)$$

where $s_{i,j}$ is the covariance between $e_{i,t}$ and $e_{j,t}$,

$$\overline{A}_i = \overline{X}_i(\overline{X}_i{}'\overline{X}_i)^{-1}\overline{X}_i{}' \qquad i = 1,\ldots,m$$
$$\overline{A}_a = \overline{X}_a(\overline{X}_a{}'\overline{X}_a)^{-1}\overline{X}_a{}'$$

and the $\overline{X}s$ are the (txp) matrices of explanatory variables in each regression.

The resulting statistics (Table 5) indicates that we cannot reject the hypothesis of perfect aggregation, at a very high confidence level.

6. Are the estimates of structural coefficients robust to aggregation?

We then evaluate the robustness of the estimated long-run parameters obtained with different methods. We follow Pesaran and Smith [1995] in comparing three different methods: i) the aggregate time series estimator; ii) the mean group estimator; iii) the pooled estimator. The use of cross-sections may also give consistent estimates of the long-run effects. Although we performed the exercise, we do not show results for the cross-section estimator, as the number of observations is too limited for any meaningful inference. However, cross-section plots (Figure 7) suggest a positive cross-country correlation between per-capita money and per-capita income (although with less than unitary elasticity) and a negative, but very imprecise, correlation between the money/GDP ratio and the interest differential.

The *aggregate time series* estimator, obtained from the equation estimated in the previous section, is shown in the first row of Table 5.

In the second row, the *mean group estimator* (MG) is obtained from (2). The long-run area coefficients are computed as:

$$a = \sum_{k=1}^{11} w_k a_k ; \quad b = \sum_{k=1}^{11} w_k b_k ; \quad g = \sum_{k=1}^{11} w_k g_k$$

The standard deviation of each coefficient is derived accordingly, considering the whole covariance matrix of individual coefficients.

We obtain two main results. Income elasticity is higher than one and almost identical to the aggregate time series estimator. On the other hand, the point estimates of the coefficients on the interest rate differentials are larger than in the previous case. However, neither of the coefficients on interest rates is statistically significant, reflecting the limited efficiency of the MG estimator.

A *pooled* estimator (third row of Table 5) is obtained by estimating the set of equations (2) as a panel with fixed effects, constraining all parameters to be the same across countries but allowing the constant to vary. The results are notably worse than in the previous case: the standard error of the estimate is larger, while the sign of the interest rate coefficients is either wrong or non-significant.

We also follow a fourth approach suggested by Pesaran, Shin and Smith [1999]: when groups are heterogeneous but there are reasons to expect some similarities, they suggest using an "intermediate" estimator between the mean group and the traditional pooled estimator. This estimator restricts only the long-run parameters (all or some of them) to be equal across countries, while allowing the intercept, the short-run coefficients and the error variances to differ. The "*pooled mean group*" (PMG) estimator (fourth row of Table 5) is based on the assumption that long-run coefficients are the same for only a subset of countries, while the dynamics are still allowed to differ. Such an approach was applied to euro-area money demand by Golinelli and Pastorello [2000],[23] using a slightly different dataset and specification.

We tested the assumption that all the long-run coefficients are the same for the eleven countries and largely rejected it (Table 6). Only for the group comprising Germany, Austria and the Benelux countries is the hypothesis of equality of each individual coefficient not

[23] They reject the hypothesis of equality for all countries and accept it only for a group of core countries, including Germany and France and excluding Italy and Spain.

rejected at the 5 per cent level. We apply the PMG estimator by imposing this constraint (last line of Table 5).[24]

The estimate of income elasticity still proves robust (although it is now not significantly different from one); both the value and the precision of the estimate of the interest elasticity are now larger.

For comparison, the area-wide dynamics resulting from the aggregate time series and the PMG estimator are shown in figure 8 and 9; these report the response of area-wide M3 to an increase in the short-term rate and in output, according to the aggregate time series and PMG estimators respectively.[25] To compute the response to the short rate, we modeled the effect of the short rate on both interest rate differentials entering the equations. A simple regression suggest that a permanent increase in the short rate has an impact of 0.6 (0.4) and an equilibrium effect of 0.4 (0.5) on the short (long) - own rate spread.

The area-wide dynamic properties are not too different across models over the relevant horizon; however, the long-run differences discussed in the preceding section show up. A permanent increase of one percentage point in the short rate has a negative effect on money demand, which reaches half of its effect after one year, almost the full effect after two years; the long-run semi-elasticity is larger for the PMG estimator. After two years the response of M3 to a one point permanent increase in output is about one percent in both models.

7. The features of national money demands

In this section we move on to a closer exam of the individual countries' characteristics that underlie the area-wide results.

As a first step we examine the individual countries' results underlying the PMG estimator commented in the previous section. The corresponding long-run coefficients are reported in table 7, together with their standard errors.

---

[24] The joint hypothesis is still rejected. However, Pesaran, Shin and Smith [1999] note that in small panels some rejections of equality of coefficients may derive from specification errors or small sample bias; they argue that pooling may still provide more accurate estimates.

[25] Luxembourg, Ireland and Finland are not included in the latter simulation.

It is evident from table 7 that some of the underlying estimates at the country level are quite imprecise. The coefficients associated with the yield differentials are not statistically significant outside the group of "core" countries; the loading coefficients on the ECM term are often small and have large standard errors.[26] The imprecise estimates may come from the imposition of the same functional form, abundantly parametrised, on each equation.

For this reason we also present two alternative sets of estimates of the national long-run parameters. Table 8 presents a set of more parsimoniously parametrised national autoregressive distributed lag models (ARDL), following Pesaran, Shin and Smith [1999]. The appropriate lags for each variable were chosen based on the Akaike criterion;[27] diagnostic tests for the individual equations are also included. Table 9 also reports the long-run coefficients estimated with the Phillips and Hansen [1990] fully modified OLS estimator, which treats the short-run dynamics in a non-parametric way.

Unsurprisingly, the results are mixed across the different methods. However, the empirical evidence allows us to draw some general conclusions. First, the performance of some national equations is not very satisfying, compared with that of the aggregate equation. Second, estimates of the income elasticity are relatively robust and not too different across countries. Third, the effect of interest rates is quite difficult to capture with precision; however, some regular patterns emerge, with money more sensitive to interest rates in one group of countries, less sensitive in a second group.

As far as the *performance of national equations* is concerned, the standard error of the estimate is around 1 percent for most countries, with the exception of Finland and Ireland (about 2 percent). The estimated coefficients on the ECM term usually have the expected, negative sign for most countries, suggesting the existence of an adjustment mechanism between money and the right-hand side variables; for Luxembourg and Ireland the

---

[26] While the negative values of the loading coefficient point to the existence of an adjustment mechanism, for most countries the hypothesis of no cointegration cannot be rejected, taking into account the critical t-values reported in Pesaran, Shin and Smith [1999b].

[27] We alternatively employed the Schwartz criterion. While it resulted in a more parsimonious choice of the number of lags, the results are qualitatively similar.

coefficient has the wrong sign.[28] The assumption of absence of cointegration is not strongly rejected in many cases. As far as stability is concerned, the relatively good performance of the aggregate equation hides national differences, which compensate each other in the aggregate time series. The one-step ahead residuals in figure 12 are outside the 95 percent band in France and Germany in 1990 and 1994, in Italy in 1996 and in Spain in 1993.

The less satisfactory performance of national equations may signal that the estimates of individual equations may be biased because of specific omitted variables or measurement errors that are correlated with the regressors. Working on individual equations it would be possible to experiment with different specifications or data until plausible estimates were obtained, but this is not possible when the functional form is constrained to be the same across individual relations (Pesaran, Shin and Smith [1999]).

The *estimates of the income elasticity* are relatively more robust across countries and methods. The long-run income elasticity is above one in most countries according to all estimators (a notable exception is Ireland). The short-run response of money to income is slightly more differentiated. Figure 11 shows the dynamic response of M3 in the different countries to an increase in GDP based on the PMG estimator.

The *interest rate elasticity* is not estimated precisely for some countries, being either non-significant or different across various methods. A common pattern emerges, however. In a group of countries money demand is interest sensitive (Germany, Austria, the Benelux countries in the PMG estimates; Germany, Belgium and Luxembourg in the parsimonious ARDL estimates; the same countries, plus Ireland and Finland, in the Phillips-Hansen estimator); in a second group the interest elasticity is always either non-significant or negligible (notably the Latin countries: Italy, France, Spain). For illustration purposes, figure 10 shows the dynamic response to an increase in interest rates, according to the PMG estimator.

[28] For these two countries, the specification bias due to cross-border holdings of monetary instruments is likely to be substantial.

The difficulty of finding a significant interest elasticity may reflect the need to model more carefully some national specificities, which should include structural breaks and a more precise definition of the rate on the alternative asset in each country. Still, the failure to find a significant negative interest elasticity in some countries is surprising, as it stands in contrast with previous estimates of money demand.

The current definition of M3, which is broader than the definition of many pre-EMU national aggregates, may account for this difference. In some countries, in the sample period the alternative to money was represented by short-term securities, which were excluded from the definition of money, whereas they are included in the current definition of M3.[29] This may explain the finding of a low elasticity to interest rates in these countries; the same need not hold for countries where long-term securities were the main substitutes for money in the sample period.

In addition, in some countries there may be a measurement problem for the opportunity cost of money, if the long-term rate is used. In the sample period movements in the long-term rate may have mostly reflected changes in risk premia, particularly in countries more exposed to the tensions in the European Monetary System in the period 1992-96. Under this condition, an increase in the long-term rate vis-a-vis the own rate of return on money does not necessarily lead to a shift from money to long-term assets (risk-adjusted rates should enter the opportunity cost of money). This situation may since have changed, as the long-term rate is almost the same in all countries since 1999; the current interest elasticity could be higher than estimates based on historical data would suggest.

8. Conclusions

We have addressed the issue of the properties of money demand in the euro area, in order to assess whether information on national contributions to monetary aggregates can improve our understanding of the behavior of euro-area money demand.

---

[29] Compared with the definition of M2 adopted in Italy before EMU, the assets now comprising M3 also include post office bills, short-term repurchase agreements and money market paper.

An aggregate area-wide money demand has satisfactory properties, confirming the conclusions of the previous literature. The assumption of perfect aggregation cannot be rejected. Aggregation or pooling of national data makes it possible to cancel out idiosyncratic instabilities of national money demands and improve the precision of the estimate of structural coefficients.

However, the estimates of the long-run coefficients of money demand are somewhat sensitive to the choice among aggregate time series and panel estimators. The estimate of the income elasticity is more robust; the estimate of the interest elasticity is more sensitive.

Over the sample period, differences in national money demand functions exist, owing to the different features of national financial structures and markets. The interest elasticity of area-wide money demand stems from a significant elasticity to interest rates in some countries and a smaller response in others. This may be explained by the fact that in some of the latter countries most substitutes of money were short-term instruments, which are now included in the broader definition of M3. Moreover, the satisfactorily stable area-wide money demand function hides a number of country-specific episodes of instability.

The national differences in interest elasticity may explain the difficulty in precisely estimating the area-wide interest elasticity, which has been common to the literature on euro-area money demand. In addition, the existence of a few country-specific structural breaks suggests that knowledge of the institutional factors behind these episodes at the national level may make it possible to judge the information content of money better. Beyond these episodes, however, our results do not suggest that the differences in the income elasticity of money demand or in the standard errors of the estimate of national equations are large. In absence of large differences, the area-wide policymaker cannot systematically exploit national information.[30]

We conclude that the area-wide equation is an appropriate analytical tool and that a systematic reference to country-specific monetary indicators is not warranted. By contrast, national information may be useful to interpret “special factors”, i. e. institutional events that

[30] See the discussion in appendix I.

may affect temporarily or permanently money demand.[31] Recently, the importance of such an approach has been emphasized by Orphanides and Porter [2000], who argue that, once the institutional information available to the monetary authority is properly considered to account for equilibrium changes in velocity, US money is still a valuable indicator. In the euro area, to a large extent, special factors must still be looked for at the national level.

[31] For an account of the role played by the analysis of “special factors” in the ECB, see Masuch, Pill and Willeke [2000].

Appendix I. A model of the information content of national contributions

Assume an IS-LM model:

(1)
$$\begin{aligned} y &= -l r + e \\ m &= ky - ar + h \end{aligned}$$

where y, m, r are respectively output, money and the interest rate, defined as deviations from equilibrium values, while e and h are shocks with zero mean and known variance. The policymaker only observes m and r. The policy rule that makes the best use of the information contained in money can be obtained setting expected y to zero, given r and m:

(2)
$$E(y|m,r) = E(e|m,r) - l r = 0$$

The optimal policy is found by substituting the least squares predictor for e, given m and r, into (2). Under the simplifying assumption of cov(e, h)= 0, this yields the optimal rule linking the interest rate to money (Poole [1970], Friedman [1990]):

(3)
$$r = \frac{ks_e^{\ 2}}{s_h^{\ 2}\, l - ks_e^{\ 2} a}\, m$$

where the weight of m depends on the relative variances of the monetary and real disturbances.

The same problem can be addressed in a two-country setting with a single monetary policy:

(4)
$$\begin{aligned} y_i &= -l_i r + e_i \\ m_i &= k_i y_i - a_i r + h_i \end{aligned}$$

where i(=1,2) stands for the country and r is common to both countries.

The solution to (4) is again obtained by setting the expected deviation of aggregate output from equilibrium to zero, conditional on the interest rate and on information on both components of money:

(5)
$$E(y_1 + y_2|m_1, m_2, r) = E(e_1 + e_2|m_1, m_2, r) - (l_1 + l_2) r = 0$$

Considering the reduced-form equations for money in country *i*, $m_i = k_i e_i - (k_i l_i + a_i)r + h_i$, and assuming for simplicity that all the cross-equation and cross-country covariances of the disturbances are zero, it turns out that:

(6) $$E(e_1 + e_2 | m_1, m_2, r) = \frac{1}{(k_1^2 s_{e1}^2 + s_{h1}^2)(k_2^2 s_{e2}^2 + s_{h2}^2)} \begin{pmatrix} m_1 + (k_1 l_1 + a_1)r \\ m_2 + (k_2 l_2 + a_2)r \end{pmatrix}^T \begin{pmatrix} k_1 s_{e1}^2 (k_2^2 s_{e2}^2 + s_{h2}^2) \\ k_2 s_{e2}^2 (k_1^2 s_{e1}^2 + s_{h1}^2) \end{pmatrix}$$

Substituting (6) into (5) and rearranging the terms, we get the optimal interest rate rule:

(7) $$r = \frac{[b k_1 s_{e1}{}^2 (m_1) + (1-b) k_2 s_{e2}{}^2 (m_2)]}{[(b s_{h1}{}^2 l_1 + (1-b) s_{h2}{}^2 l_2)] - [(b k_1 s_{e1}{}^2 a_1 + (1-b) k_2 s_{e2}{}^2 a_2]}$$

where:

$$b = \frac{k_2^2 s_{e2}^2 + s_{h2}^2}{k_1^2 s_{e1}^2 + s_{h1}^2 + k_2^2 s_{e2}^2 + s_{h2}^2}$$

The national components of money, m_1 and m_2, enter (7) with different coefficients only when there are large cross-country differences in either the variance of real and monetary disturbances or in output elasticity. In this case, more weight should be assigned to the monetary contribution of the country characterized by smaller variance of monetary disturbances or by larger variance of real disturbances (the latter result conforms to the intuition of the original Poole model).[32] When the differences across countries are small, there is little gain from considering the national contributions rather than the aggregate m. If the variance of the real and monetary disturbances and output elasticity are equal ($s_{e1}^2 = s_{e2}^2 = s_e^2$, $s_{h1}^2 = s_{h2}^2 = s_h^2$, $k_1 = k_2 = k$), (7) becomes the analogue of (3):

(8) $$r = \frac{k s_e{}^2}{s_h{}^2 (l_1 + l_2) - k s_e{}^2 (a_1 + a_2)} (m_1 + m_2)$$

[32] By contrast, the national components of money do not enter (8) with different coefficients when there are only differences in interest rate elasticity.

Appendix II. The dataset[33]

The data employed in the paper cover the period 1982-1999.

Money M3 is the non-seasonally adjusted definition adopted by the Eurosystem. M3 is composed of currency in circulation and other liabilities issued by monetary and financial institutions (MFIs) held by euro-area residents other than the Central Government and MFIs. These liabilities comprise overnight deposits, deposits with agreed maturity up to 2 years, deposits redeemable at notice up to three months, repurchase agreements, money market fund shares, money market paper, debt securities with maturity up to two years. National contributions to M3 amount exactly to M3. They may be reconstructed aggregating the relevant items of the balance sheet of the MFIs at the national level. Data for the period before 1997 are reconstructed based on not fully harmonized national data.[34] All data are converted into euro by applying the irrevocable conversion rates fixed on 31 December 1998. The publications of some central banks include recent data on national contributions.

The German series has a break in June 1990 due to German unification; we corrected it by reconstructing the new series on the basis of the month-to-month growth rates of the old one (the same adjustment is then applied to euro-area M3).

Quarterly GDP at 1995 prices is from Eurostat (based on ESA95 criteria where available). For Ireland, Luxembourg, Austria and Finland the quarterly series are those constructed by Golinelli and Pastorello [2000], disaggregating the annual series using industrial production. All data were converted into euros by applying the irrevocable conversion rates fixed on 31 December 1998.

GDP weights used in the construction of area-wide variables are based on the 1994-99 average, at PPP exchange rates.

---

[33] This appendix was prepared by Claudio Trevisan.

[34] Italian data for M3 contributions can be found in Bank of Italy [2000]. A different, unofficial reconstruction of national contributions to M3 based on publicly available sources was made by Golinelli and Pastorello [2000].

The *price level* is measured by the national CPI indexes (the national HICPs, harmonized indices of consumer prices, were not available for the whole period).

For the *long-term interest rate*, we used the government 10-year benchmark security adopted for the Maastricht convergence criteria for all the countries. For the earlier years, when these are not available, we used the government bond yield (from IMF *International Financial Statistics*). For Portugal, the series from 1980 to 1990 is the nominal rate on government bonds, before tax, taken from Marques and Lopes [1992].

From 1999, the euro-area *short-term interest rate* is *Euribor*. Before 1999, the euro-area rate is calculated on the basis of national rates weighted by GDP shares.

National short-term interest rates are 3-month interbank rates. Sometimes data from different sources had to be used for earlier periods. For Italy, we used the 3-month T-bill rate before 1990. For Finland, we used the call money credit rate until 1986, then the 3-month CDs rate. For Portugal, it was not possible to identify a satisfactory measure of the short-term rate for the 1980s. Only to maintain the functional form adopted for the other countries, for this period we subtracted the spread between the long and short rates included in IMF *International Financial Statistics* from the long rate defined above.

The construction of the *own rate of return on M3*, both for the euro area and national contributions, is based on the rates of return on three components: currency (equal to zero), marketable instruments (whose rate was assumed to be equal to the short-term market rate) and bank deposits. Where necessary, a further distinction is made between overnight deposits, deposits with agreed maturity up to 2 years and deposits redeemable at notice up to 3 months.

The interest rates on bank deposits (either total or individual components) are obtained, whenever possible, from comparable data as currently defined by the national central banks. Where these were not available, we have used national data on interest rates on similar bank liabilities. However, the information available varied greatly for different countries, categories and periods. Accordingly, whenever possible we used data for the corresponding category of deposits from domestic sources, although these are not harmonized; a constant adjustment was used in case of a level break. When the former information was not available, we used as a proxy the yield of a similar instrument (for example, a different category of deposit).

All rates of return are then weighted with the share of the corresponding components in national contributions to M3. The weights are available monthly for the period 1998-1999. For the period from 1980 to 1997 we use constant weights, based on the composition of national contributions in the fourth quarter of 1997, obtained from the elementary data for monetary and financial institutions' balance sheets. The area-wide own rate of return on money was then computed as a weighted average of the national ones, using M3 contributions as weights. We compared the results for the own rate with the series that can be constructed, with the same methodology, using directly the area-wide yields published by the ECB; we found no major discrepancies.

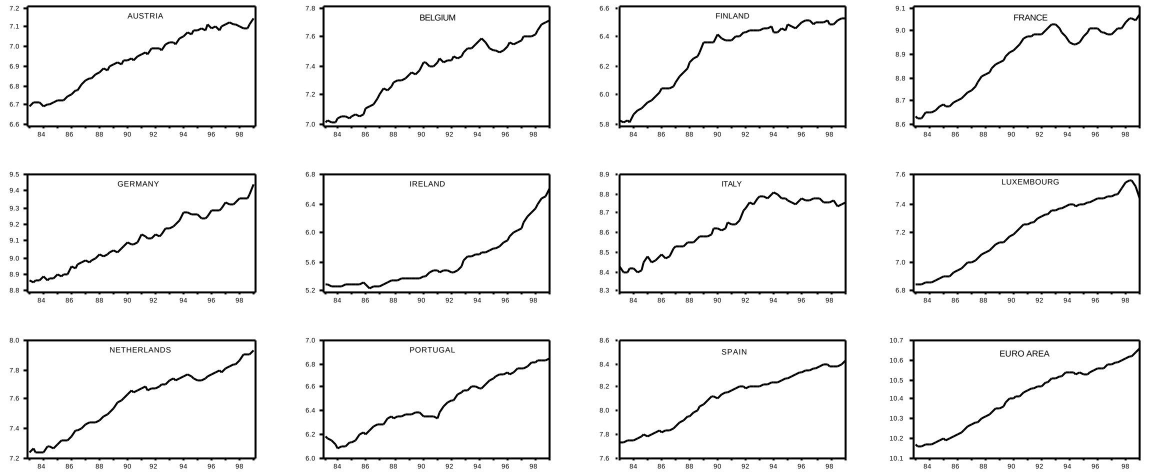
We compared our area-wide variables, constructed by aggregating national data, with the dataset included in Brand and Cassola [2000]. The differences for nominal money, GDP and long and short rates are negligible; however, they use the GDP deflator, instead of the consumer price index, as a measure of the price level.

Tables and figures

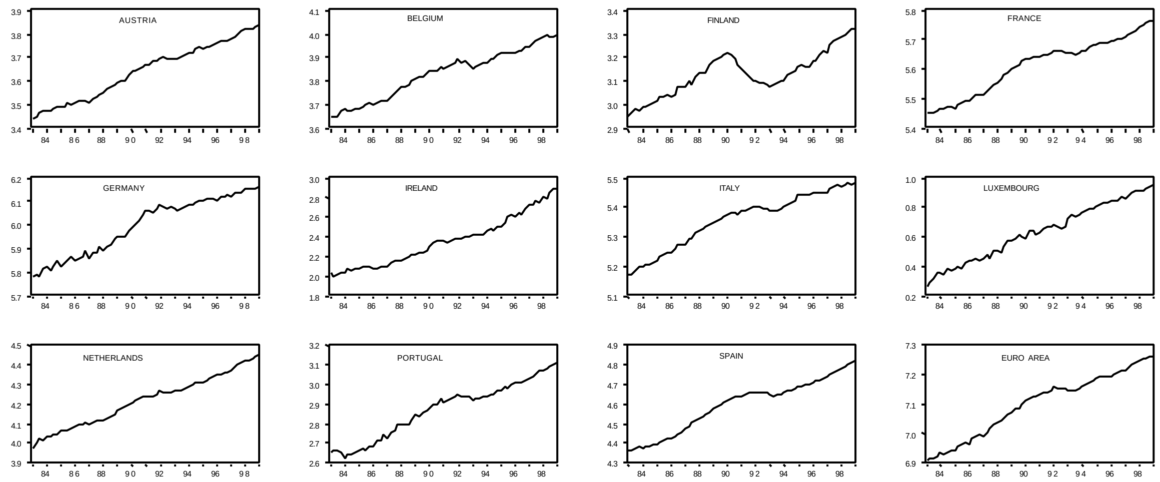
Figure 1

REAL M3 AND GDP

Real M3



GDP



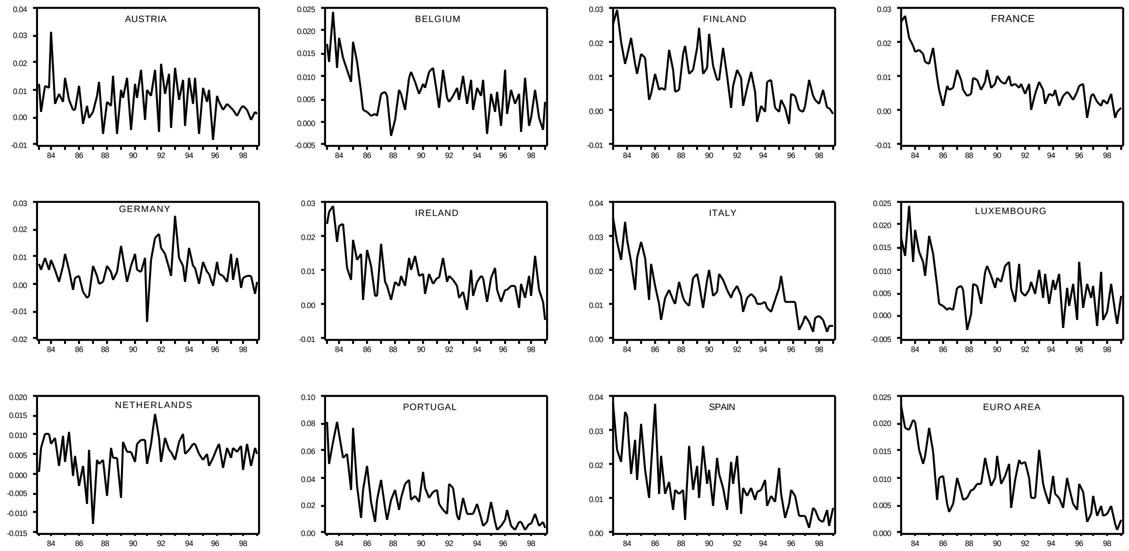
Real M3: log- difference between money (M3 and national contributions, not seasonally adjusted) and CPI (respectively euro-area and national indices). GDP: log of GDP at 1995 prices. See Appendix II for details.

.

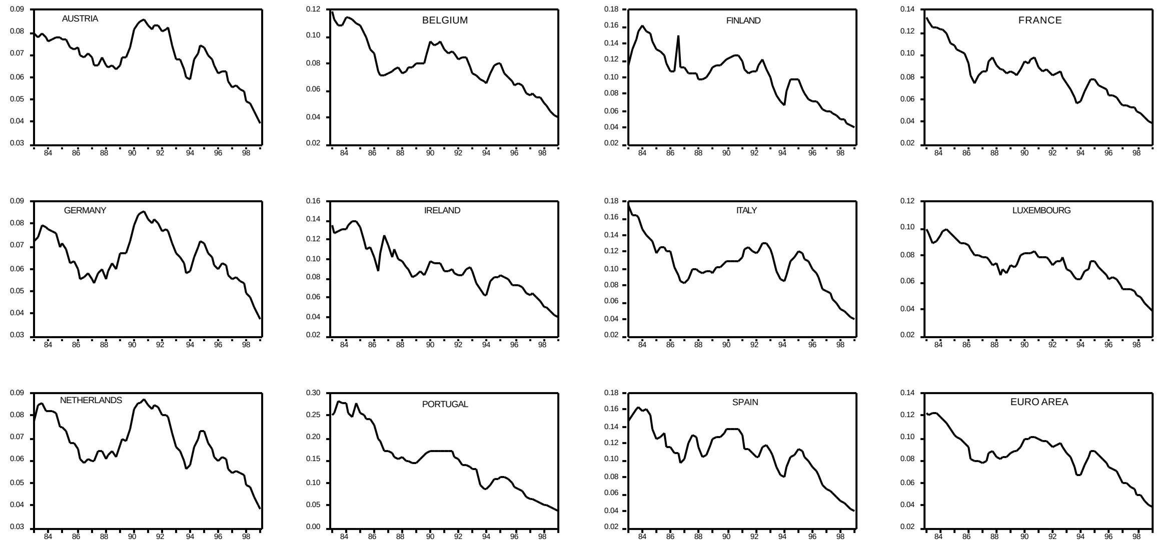
Figure 2

PRICES AND LONG-TERM INTEREST RATES

Prices (quarterly rates of change)



Long-term interest rate

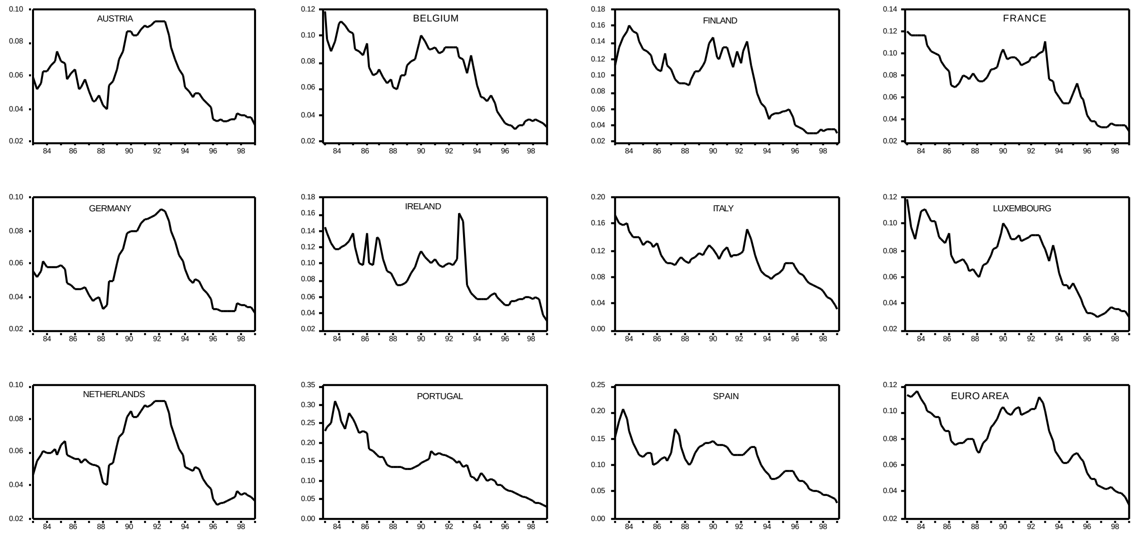


Prices: quarterly log-difference of the CPI, not seasonally adjusted. Long term rates: government bond yield (10-year benchmark when available).

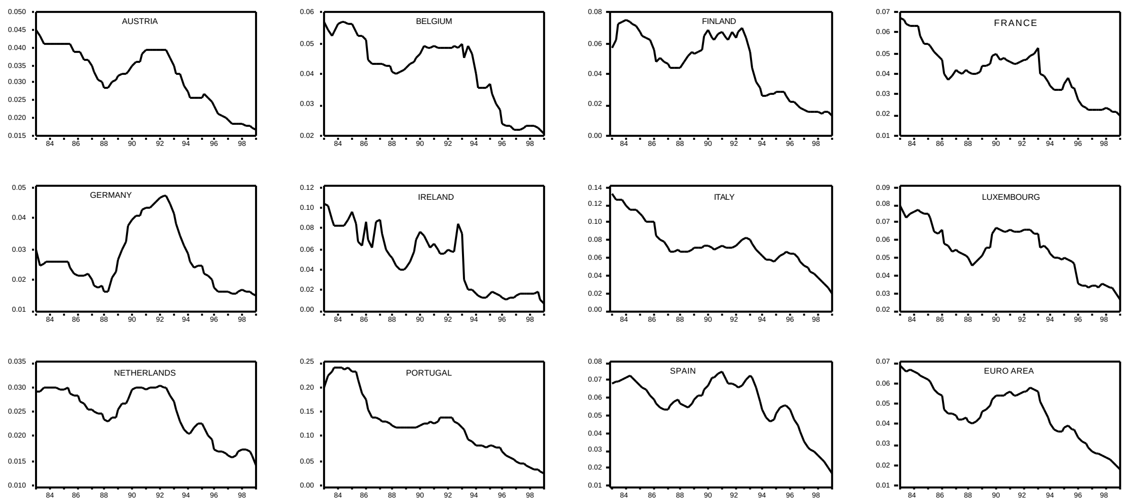
Figure 3

SHORT-TERM INTEREST RATES AND RATE OF RETURN ON M3

Short-term interest rate



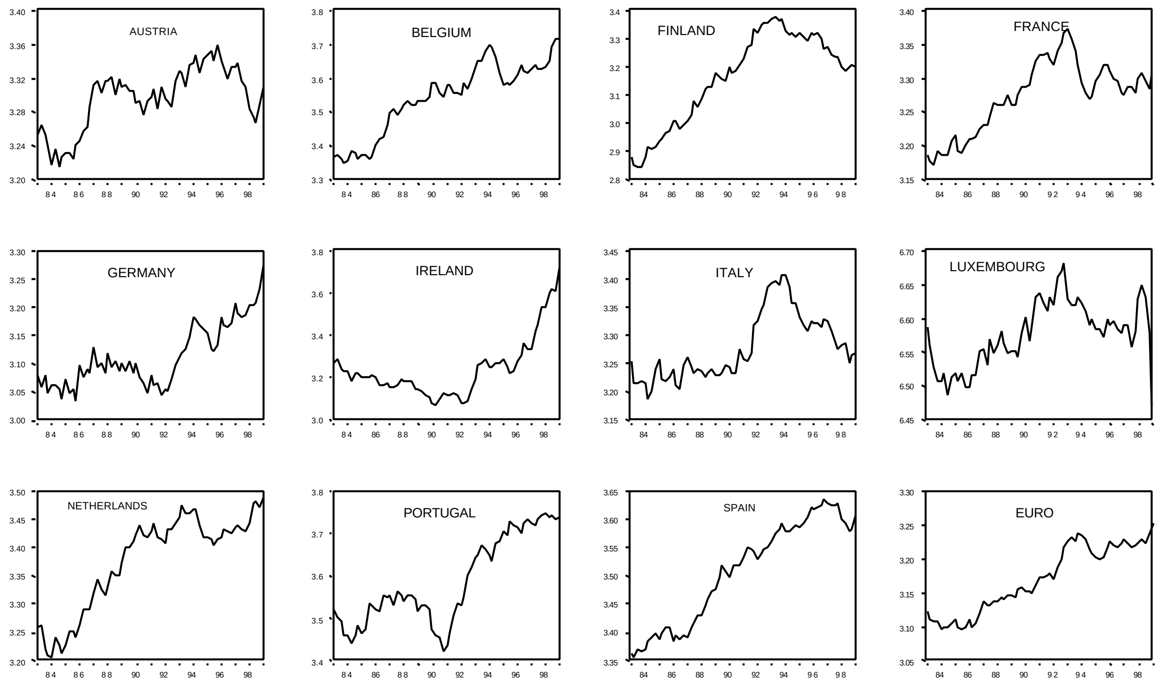
Rate of return on M3



Short-term rates: three-month interest rates, national sources. Rate of return on M3: weighted average of the yields on bank deposits, marketable instruments, currency (the latter yield equal to zero). Authors' calculations based on national sources. See appendix II for details.

Figure 4

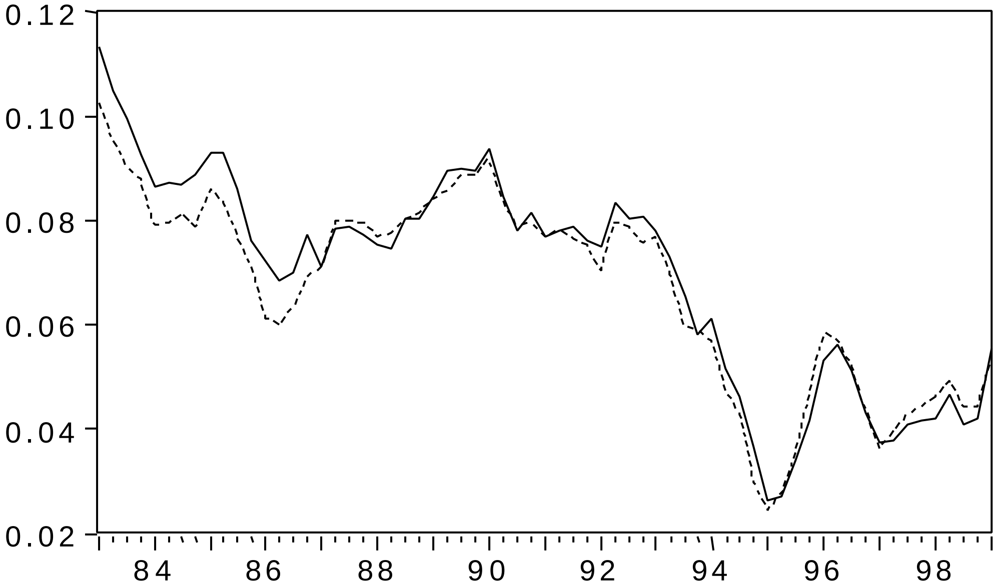
MONEY/GDP RATIO



Log difference between real money and GDP. National sources. See appendix II for details.

Figure 5

M3 GROWTH
(4-quarter log difference)

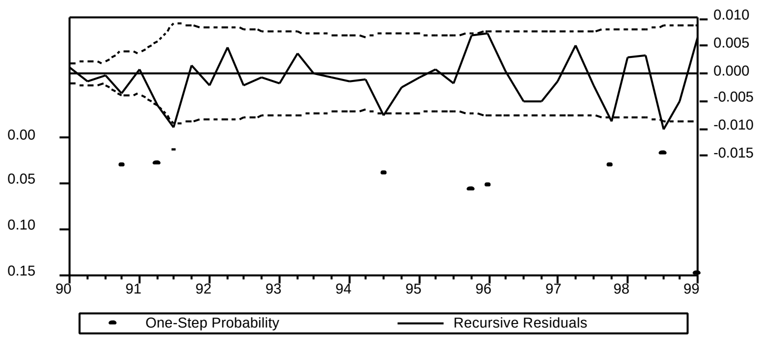


Dotted line: 4-quarter difference of log(M). *Solid line*: 4-quarter difference of $\Sigma w_i \log(M_{it})$, where M_i is the contribution of country i to area-wide money and w_i is the GDP weight of country i.

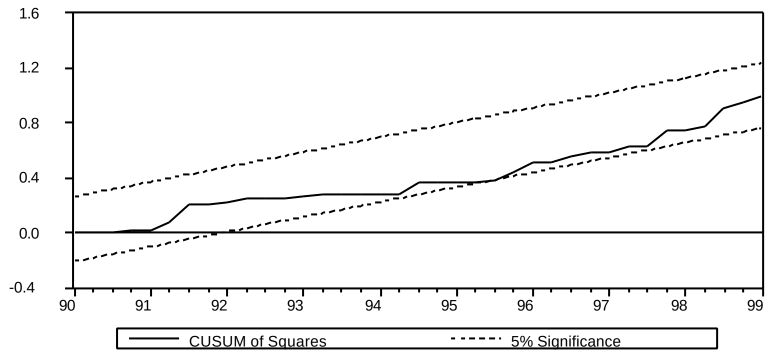
Figure 6

AGGREGATE TIME SERIES EQUATION: STABILITY TESTS

One-step ahead prediction errors



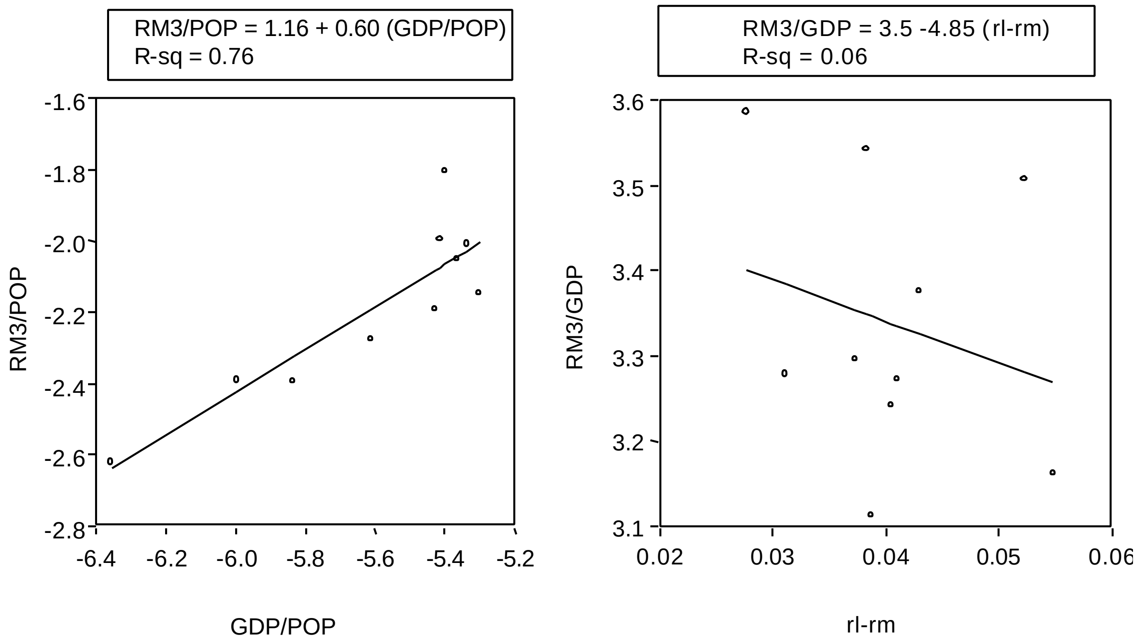
Cusum-sq tes



Stability test conducted on the equation in Table 3.

Figure 7

CROSS-SECTION PLOTS

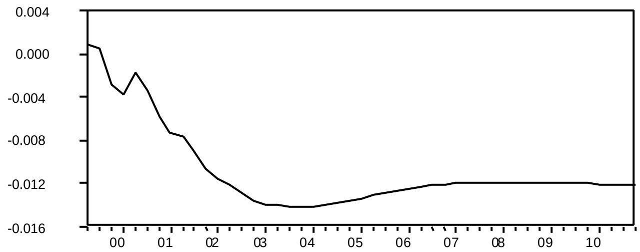


RM3/GDP: ratio between real M3 and GDP. M3/POP: per capita real M3. GDP/POP: per capita GDP. Logs of 1980-1999 averages. Luxembourg is excluded.

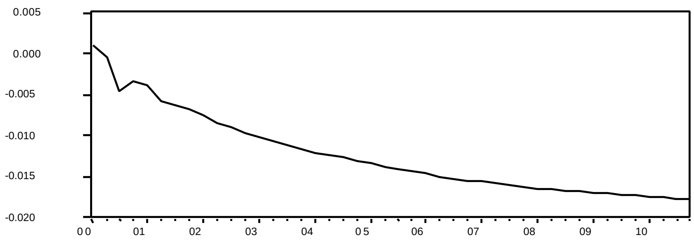
Figure 8

RESPONSES OF M3 TO INTEREST RATES

Aggregate time series estimator



Pooled Mean Group estimator

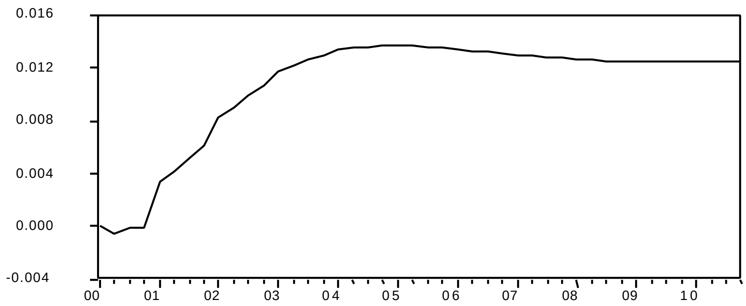


Response of the log of real M3 to a permanent increase of one percentage point in the short rate.

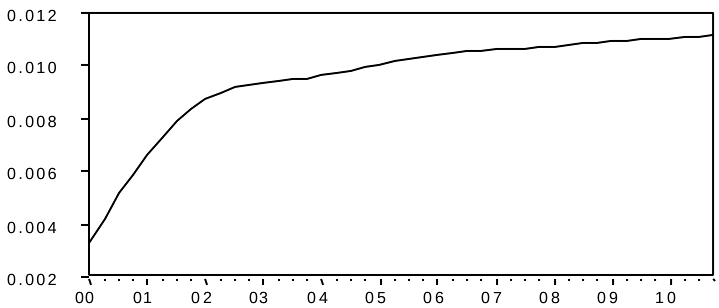
Figure 9

RESPONSES OF M3 TO AN INCREASE IN GDP

Aggregate time series estimator



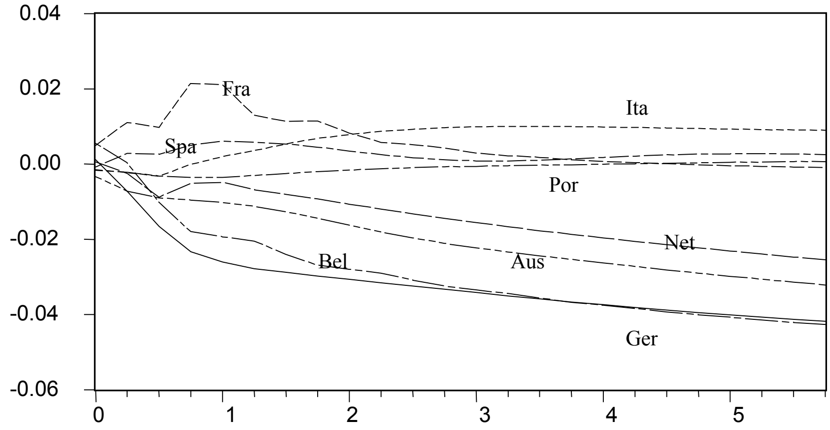
Pooled Mean Group estimator



Response of the log of real M3 to permanent increase of one percent in GDP.

Figure 10

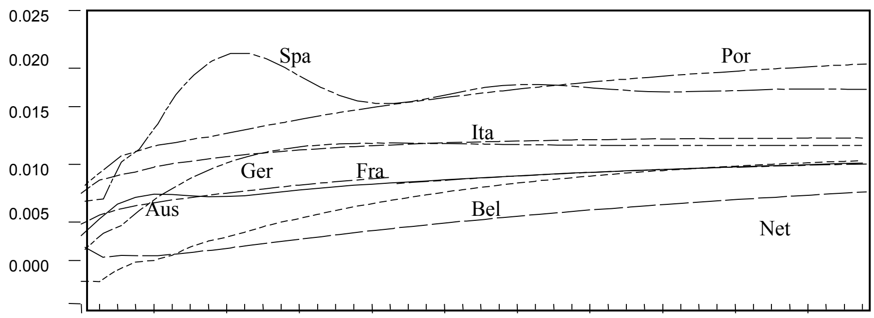
RESPONSES OF NATIONAL CONTRIBUTIONS TO M3 TO INTEREST RATES



Response of the log of national contributions to M3 (in real terms) to a permanent increase of one percentage point in the short rate (see text).

Figure 11

RESPONSES OF NATIONAL CONTRIBUTIONS TO M3 TO GDP

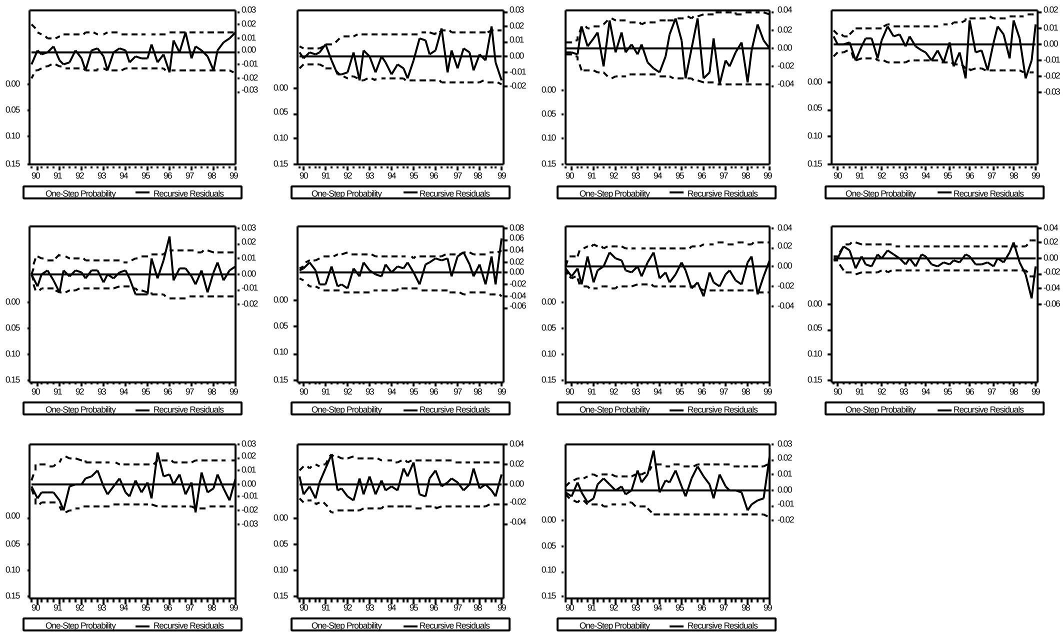


Response of the log of national contributions to M3 (in real terms) to a permanent increase of one percent in GDP (see text).

Figure 12

STABILITY TESTS

One-step ahead prediction errors



Stability tests conducted on national equations. The graphs refer to: Austria, Belgium, Finland, France (first row); Germany, Ireland, Italy, Luxembourg (second row); Netherlands, Portugal, Spain (third row).

Table 1

UNIT ROOT TESTS

	(m-p)		(rs-rm)		(rl-rm)	
	level	Δ	level	Δ	level	Δ
Austria	-2,10	-4,96**	-2,07	-3,56**	-1,89	-4,52**
Belgium	-2,39	-3,82**	-2,08	-5,09**	-2,11	-4,16**
Finland	-0,53	-4,88**	-1,08	-4,00**	-1,41	-4,30**
France	-1,55	-3,66**	-0,82	-4,10**	-1,56	-4,51**
Germany	-1,37	-4,90**	-0,87	-3,63**	-1,92	-3,70**
Ireland	1,15	-4,03**	-2,55	-4,94**	-2,82	-3,28**
Italy	-1,14	-3,77**	-1,75	-4,54**	-2,00	-3,84**
Luxembourg	0,69	-2,61	-2,02	-5,18**	-1,67	-4,34**
Netherlands	-1,52	-4,24**	-1,44	-3,27**	-1,60	-3,56**
Portugal	-2,70	-4,57**	-2,02	-4,39**	-2,57	-3,73**
Spain	-1,67	-4,30**	-1,38	-5,44**	-1,28	-5,18**
Euro area	-1,99	-3,73**	-1,49	-4,86**	-3,14	-5,68**

	y		π	
	level	Δ	level	Δ
Austria	-2,49	-7,43**	-4,99**	-12,4**
Belgium	-3,01	-4,65**	-2,08	-8,78**
Finland	-2,22	-16,1**	-3,17	-8,95**
France	-1,40	-4,69**	-1,94	-6,25**
Germany	-1,57	-4,92**	-3,30**	-9,10**
Ireland	-0,80	-5,15**	-6,14**	-8,07**
Italy	-1,90	-3,65**	-2,28	-7,55**
Luxembourg	-3,06	-6,27**	-2,68	-4,46**
Netherlands	-2,77	-4,56**	-2,26	-8,43**
Portugal	-1,79	-3,22**	-3,52**	-6,42**
Spain	-2,60	-2,56	-3,09	-9,80**
Euro area	-2,25	-4,07**	-2,06	-8,78**

ADF Statistics. ** Indicates rejection of the hypothesis that the series is I(1).

Table 2

JOAHNSSEN COINTEGRATION TEST

	r=0	r=1	r=2	r=3	r=4	Number of cointegrating vectors at 1%
M/P, Y	17.11*	1.21	-	-	-	none
M/P, rs-rm	2.6	0.06	-	-	-	none
M/P, rl-rm	0.19	0.01	-	-	-	none
M/P,Y,rs-rm	23.2	4.83	0.03	-	-	none
M/P, Y, rl-rm	42.2**	12.8	2.2	-	-	1
M/P,Y,rl-rm,rs-rm	55.8**	22.1	6.2	0.1	-	1
M/P, π	12.9	0.6	-	-	-	none
M/P, Y, π	34.1*	10.2	3.4	-	-	none
M/P, rs-rm, π	28.0	3.3	0.2	-	-	none
M/P, rl-rm, π	25.0	11.9	0.6	-	-	none
M/P,Y,rs-rm, π	48.8*	16.9	6.4	0.1	-	none
M/P, Y, rl-rm, π	55.8**	24.5	11.8	5.1	-	1
M/P,Y,rl-rm,rs-rm, π	82.4**	41.7	19.9	6.3	0.1	1

* Indicates indicates rejection at 5%. ** Indicates rejection at 1%.

Table 3

MONEY DEMAND IN THE EURO AREA

	coefficient	Std. Error
cost.	0.20	0.04
ϕ	-0.12	0.03
y_{t-1}	1.26	0.06
$(rl-rm)_{t-1}$	-3.36	1.07
$(rs-rm)_{t-1}$	-0.08	0.43
$\Delta \log (M/P)_{t-4}$	0.35	0.10
$\Delta \log(Y)_{t-1}$	-0.22	0.09
$\Delta \log(Y)_{t-4}$	0.23	0.09
$\Delta(rl-rm)_{t-4}$	0.48	0.16
$\Delta(\pi)_t$	-0.25	0.15
$\Delta(\pi)_{t-1}$	-0.31	0.16
$\Delta(\pi)_{t-4}$	0.50	0.16
seasonal 1	-0.005	0.001
seasonal 2	-0.006	0.002
seasonal 3	-0.008	0.001

Estimated equation:

$$\Delta \log(\frac{M}{P})_t = a + j\,[\log(\frac{M}{P})_{t-1} - a\,\log(Y_{t-1}) - b\,(rl-rm)_{t-1} - g\,(rs-rm)_{t-1}] +$$

$$+ b_4 \Delta \log(\frac{M}{P})_{t-4} + c_1 \Delta \log(Y)_{t-1} + c_4 \Delta \log(Y)_{t-4} + d_4 \Delta (rl-rm)_{t-4} + f_0 \Delta p_{t-0} + f_i \Delta p_{t-1} + f_i \Delta p_{t-4}$$

Sample: 1983:3 1999:1

R-squared	0.839053	Mean dependent var	0.007828
Adjusted R-squared	0.792110	S.D. dependent var	0.007384
S.E. of regression	0.003367	Akaike info criterion	-8.345445
Sum squared resid	0.000544	Schwarz criterion	-7.835175
Log likelihood	277.8815	Durbin-Watson stat	1.950060

Breusch-Godfrey Serial Correlation LM Test:

F-statistic	0.014060	Probability	0.986043
Obs*R-squared	0.038488	Probability	0.980940

ARCH Test:

F-statistic	1.824037	Probability	0.137598
Obs*R-squared	7.022835	Probability	0.134686

Jarque-Bera Normality Test:

	0.846030	Probability	0.655069

Table 4

M3 DEMAND:
A COMPARISON OF LONG-RUN COEFFICIENTS

	y	rs	rl	rs – rm	rl-rs	π
Single equation	1.26**			-0.08	-3.36**	
	(0.06)			(0.43)	(1.06)	
Johanssen procedure	1.38**			0.41	-1.71**	
	(0.02)			(0.30)	(0.49)	
Coenen-Vega (1999)	1.14**	0.82**	-0.82**	-	-	-1.46**
	(0.06)	(0.35)	(0.35)			(.32)
Brand-Cassola (2000)	1.33**	-	-1.61**			
	(0.03)		(0.01)			
Golinelli.-Pastorello. (2000)	1.37**	-	-0.68**	-	-	-
	(0.05)		(0.31)			

Standard errors are in parentheses.

Table 5

M3 DEMAND: ALTERNATIVE ESTIMATORS OF LONG-RUN COEFFICIENTS

Estimator:	*y*	*rl – rm*	*rs - rm*	*SEE(%)*
Aggregate time-series (1)	1.26** (.06)	-3.36** (1.1)	-0.08 (0.43)	0.34
Mean Group (2)	1.25** (.11)	-3.51 (3.06)	-1.41 (1.30)	0.33
Pooled (3)	1.64** (.15)	-0.01 (1.84)	0.62 (1.40)	0.46
Pooled Mean Group (4)	1.16** (0.10)	-4.12** (1.21)	-1.27 (0.97)	0.34

Grunfeld-Griliches prediction criterion: GG = $s^2_d - s^2_a = -0.01$

Test of perfect aggregation: $\chi^2_{64} = 1.83$ [99%]

(1) See Table 4. (2) Area coefficients are obtained as a weighted average of national coefficients (GDP weights). National equations include four lags of each variable and of the first difference of inflation, plus seasonal dummies. (3) Panel estimation. All coefficients constrained to be equal across all countries. Country effects and country specific seasonal dummies included. (4) Area coefficients are obtained as a weighted average of country coefficients (GDP weights). Long-run coefficients constrained to be equal across 5 countries (Germany, Austria, Belgium, Netherlands, Luxembourg). Fixed effects included.

Table 6

TEST OF EQUALITY OF LONG-RUN COEFFICIENTS

	y	*rl - rm*	*rs – rm*	*All*
H(11)	χ^2 = 29.58 [0.00%]**	χ^2 = 31.58 [0.00%]**	χ^2 = 11.00 [35.72%]*	χ^2 = 137.81 [0.00%]**
H(5)	χ^2 = 9.02 [6.0%]	χ^2 = 2.19 [70.0%]	χ^2 = 7.69 [10.4%]	χ^2 = 32.37 [0.12%]**

Wald test. H(11): the long-run coefficients are the same for the 11 euro-area countries. H(5): the long-run coefficients are the same for Germany, Luxembourg, Belgium, the Netherlands, Austria.

Table 7

LONG-RUN COEFFICIENTS: *POOLED MEAN GROUP* ESTIMATOR

	Y	rl-rm	rs-rm	loading	SEE(%)
Aus	1.04**	-10.99**	-3.6**	-0.02**	0.79
	0.1	*2.42*	*1.41*	*0.01*	
Bel	1.04**	-10.99**	-3.6**	-0.04**	0.9
	0.1	*2.42*	*1.41*	*0.02*	
Fin	1.56	2.91	9.82	-0.04**	1.94
	0.91	*8.23*	*10.91*	*0.02*	
Fra	1.17**	-0.66	0.05	-0.09**	0.91
	0.36	*1.96*	*1.99*	*0.06*	
Ger	1.04**	-10.99**	-3.6**	-0.04**	0.74
	0.1	*2.42*	*1.41*	*0.02*	
Irl	-1.4	11.18	11.1	0.03	2.08
	2.37	*10.09*	*21.46*	*0.02*	
Ita	1.15**	3.13	-0.73	-0.09**	1.29
	0.19	*1.64*	*2.7*	*0.04*	
Lux	1.04**	-10.99**	-3.6**	0.1	1.27
	0.1	*2.42*	*1.41*	*0.04*	
Net	1.04**	-10.99**	-3.6**	-0.01	0.9
	0.1	*2.42*	*1.41*	*0.01*	
Por	2.17**	-8.42	9.13	-0.02	1.1
	0.84	*9.17*	*12.72*	*0.02*	
Spa	1.71**	1.38	-0.77**	-0.25**	0.89
	0.05	*0.48*	*0.3*	*0.05*	

National estimates underlying the Pooled Mean Group Estimator in Table 5. Standard errors are indicated below each coefficient. ** Indicates an estimate not different from zero at 5%.

Table 8

LONG-RUN COEFFICIENTS: *AUTOREGRESSIVE DISTRIBUTED LAG* MODELS

	y	rl-rm	rs-rm	loading		SEE(%)	SC	ff	norm	het
Aus	1.09**	-8.38	0.67	-0.03	ARDL(2,0,0,1)	0.6	13.90%	3%	1.20%	4.30%
	0.3	*14.23*	*2.76*	*0.04*						
Bel	1.69**	-5.41**	0.36	-0.21**	ARDL(4,4,1,2)	0.9	77.20%	12%	42.70%	6.10%
	0.09	*1.19*	*0.61*	*0.04*						
Fin	0.31	-7.72	5.35	-0.05**	ARDL(1,1,0,0)	1.8	76.80%	14.20%	80.10%	22.00%
	0.94	*7*	*6.12*	*0.02*						
Fra	1.25**	-1.15	0.58	-0.12	ARDL(4,2,4,4)	0.9	26.70%	0.40%	70.50%	19.70%
	0.31	*2.13*	*1.57*	*0.08*						
Ger	1.23**	-5.6	-4.5**	-0.08	ARDL(3,3,2,3)	0.7	19.80%	54.50%	99.10%	18.40%
	0.14	*5.4*	*1.9*	*0.06*						
Irl	-0.14	5.69	-6.01	0.03	ARDL(4,1,1,1	1.8	2.30%	0.10%	62.20%	14.70%
	1.35	*6.7*	*11.2*	*0.03*						
Ita	0.64	0.12	9.35	-0.05	ARDL(1,0,0,3)	1.21	35.00%	67.90%	96.90%	34.50%
	0.65	*3.91*	*9.16*	*0.04*						
Lux	1.11**	-14.16**	-4.77	0.06	ARDL(4,1,2,0)	1.1	16.40%	7.80%	18.10%	23.70%
	0.28	*9.55*	*5.51*	*0.04*						
Net	1.48**	-9.76	3.86	-0.07**	ARDL(2,1,0,1)	0.8	89.20%	3.90%	69.70%	23.90%
	0.17	*6.49*	*1.83*	*0.03*						
Por	2.65	-5.15	16.48	-0.01	ARDL(4,1,0,1)	1.1	27.20%	88.00%	5.60%	81.10%
	2.4	*12.9*	*36.44*	*0.02*						
Spa	1.75**	1.44	-0.6**	-0.26**	ARDL(2,4,3,3)	0.8	39.60%	76.00%	60.50%	2.20%
	0.07	*0.61*	*0.38*	*0.07*						

ARDL(*a*, *b*, *c*, *d*) stands for a model including *a* lags of real money, *b* lags of *y*, *c* lags of *rl-rm*, *d* lags of *rs-rm* (see Pesaran, Shin and Smith [1999]). The number of lags included for each variable was selected according to an Akaike criterion. *SC*: confidence level for serial correlation test; *ff*: confidence level for functional form test; *norm*: confidence level for normality test; *hex*: confidence level for heteroskedasticity test. ** Indicates that the coefficient is not statistically different from zero.

Table 9

LONG-RUN COEFFICIENTS:
PHILLIPS-HANSEN *FULLY MODIFIED LEAST SQUARES*

	y	rl-rm	rs-rm
Aus	1.25**	1.31	-0.66
	0.04	*0.9*	*0.37*
Bel	1.87**	-3.14**	0.88
	0.06	*0.51*	*0.41*
Fin	1.29**	-1.95	-4.1**
	0.34	*2.22*	*1.44*
Fra	1.57**	-0.94	2.23
	0.07	*0.54*	*0.31*
Ger	1.28**	-1.61**	-3.39**
	0.03	*0.59*	*0.41*
Irl	1.43**	1.82	-5.21**
	0.08	*1.47*	*2.62*
Ita	1.18**	3.17	-2.47**
	0.06	*0.88*	*0.97*
Lux	1.12**	-3.34**	0.79
	0.08	*2.2*	*1.4*
Ola	1.63**	-1.44**	2.24**
	0.08	*0.6*	*0.31*
Por	1.45**	-1.01	-1.85
	0.1	*1.2*	*1.2*
Spa	1.61**	0.8	-0.9**
	0.06	*0.48*	*0.29*

See Phillips and Hansen [1990]. Standard errors below each coefficient. ** Indicates that the coefficient is not statistically different from zero.

References

Ando, A. and K. Shell (1975), "Demand for money in a general portfolio model in the presence of an asset that dominates money", in Fromm, G. and L. R. Klein (eds.), *The Brookings Model: Perspective and Recent Developments*, Amsterdam, North-Holland, pp. 560-63.

Angelini, P., P. Del Giovane, S. Siviero and D. Terlizzese (2000), "Monetary policy in a multi-country monetary union: should information at the national level play a role?", Banca d'Italia, mimeo.

Angeloni, I., C. Cottarelli and A. Levy (1992), "Cross-border deposits and monetary aggregates in the transition to EMU", Banca d'Italia, *Temi di Discussione*, No. 163.

Arnold, I. J. M. (1994), "The myth of a stable european money demand", *Open Economies Review*, Vol. 5, No. 3, pp.249-59.

Calza, A., D. Gerdesmeier and Levy, J. (2001), "Euro area money demand: measuring the opportunity costs appropriately", European Central Bank, mimeo.

Coenen, G. and J. L. Vega (1999), "The demand for M3 in the euro area", European Central Bank, Working Paper, No. 6.

Baba, Y., D. Hendry and R. Starr (1992), "The demand for M1 in the USA, 1960-88", *Review of Economic Studies*, Vol. 59, No. 1, pp. 51-61.

Bank of Italy (2000), "Aggregati monetari e creditizi dell'area dell'euro: le componenti italiane. Note metodologiche e informazioni statistiche", *Supplementi al Bollettino Statistico*, n. 33 (downloadable at www.bancaditalia.it).

Brand, C. and N. Cassola (2000), "A money demand system for euro area M3", European Central Bank, Working Paper, No. 39.

Davidson, J. (1998), "Structural relations, cointegration and identification: some simple results and their application", *Journal of Econometrics*, Vol. 87, pp. 87-113.

De Grauwe, P. (2000), "Monetary policies in the presence of asymmetries", CEPR Working Paper, No. 2393.

European Central Bank (1999a), "The stability-oriented monetary policy strategy of the Eurosystem ", *Monthly Bulletin,* January.

_______________ (1999b), "The review of the reference value for money growth", *Monthly Bulletin,* December.

_______________ (2000a), "Staff Economic Projections for the Euro Area", *Monthly Bulletin,* December.

_______________ (2000b), "Annual review of the reference value for monetary growth", *Monthly Bulletin,* December.

Fagan, G. and J. Henry (1998), “Long run money demand in the EU: evidence from area-wide aggregates”, *Empirical Economics*, Vol. 23 pp. 483-506.

Fase, M. G. and C. C. A. Winder (1998), “Wealth and the demand for money in the European Union”, *Empirical Economics*, Vol. 23, pp. 507-24.

Friedman, B. (1990), “Targets and instruments of monetary policy”, in Friedman, B. and F. Hahn (eds.), *Handbook of Monetary Economics*, Amsterdam, North Holland.

Filosa, R. (1995), “Money demand stability and currency substitution in six European countries”, BIS Working paper No. 30.

Golinelli, R. and S. Pastorello (2000), “Modeling the demand for M3 in the Euro area”, manuscript (downloadable at www.spbo.unibo.it/pais/golinelli/ GoPaENG_ref.pdf).

Grunfeld, Y. and Z. Griliches (1960), “Is aggregation necessarily bad?”, *The Review of Economics and Statistics*, Vol. 17, pp.1-13.

Marques, R. C. and C. Lopes (1992), “Choosing an aggregate for monetary policy: a cointegration approach”, Banco de Portugal, Research and Statistics Department, Working Paper, No. 24-92.

Masuch, K., H. Pill and C. Willeke (2000), “Framework and tools of monetary analysis”, paper presented at the Seminar on “Monetary Analysis: Tools and Applications” held at the European Central Bank, November (downloadable at www.ecb.org).

Monticelli, C. and L. Papi (1996), *European integration, monetary co-ordination and the demand for money*, Oxford, Clarendon Press.

Mulligan, C. B. and X. Sala-i-Martin (1992), “U.S. money demand: surprising cross-sectional estimates”, *Brookings Papers on Economic Activity*, Vol. 2.

OECD (2000), “Could aggregation be misleading?”, in OECD, *Emu one year on*, Paris.

Orphanides, A. and R. Porter (2000), “Money and inflation: the role of information regarding the determinants of M2 behavior”, paper presented at the Seminar on “Monetary Analysis: Tools and Applications” held at the European Central Bank, November (downloadable at www.ecb.org).

Pesaran, M. H., R. G. Pierse and M. S. Kumar (1989), “Econometric analysis of aggregation in the context of linear prediction models”, *Econometrica*, No. 4.

Pesaran, M. H., Y. Shin and R. P. Smith (1999), “Pooled mean group estimation in dynamic heterogeneous panels”, *Journal of the American Statistical Association*, Vol. 94, pp. 621-34.

Pesaran, M. H., Y. Shin and R. P. Smith (1999b), “Bounds testing approaches to the analysis of long-run relationships ”, University of Cambridge, Dept. of Applied Economics, Working Paper, No. 7.

Pesaran, M. H. and R. Smith (1995), “Estimating long-run relationships from dynamic heterogeneous panels”, *Journal of Econometrics*, Vol. 68, pp. 79-113.

Phillips, P. C. B. and B. E. Hansen (1990), “Statistical inference in instrumental variables regression with I(1) processes”, *Review of Economic Studies*, Vol. 57, No. 189, pp. 99-125.

Poole, W. (1970), “Optimal choice of monetary policy instruments in a simple stochastic macro model”, *Quarterly Journal of Economics*, Vol. 84.

Trecroci, C. and J. L. Vega (2000), “The Information content of M3 for future inflation in the euro area”, European Central Bank, Working Paper, No. 33.

Vega, J. L. (1998), “The ALP Long-Run Demand Function” in Malo De Molina, J. L., J. Vinals and F. Gutierrez (eds.), *Monetary Policy and Inflation in Spain*, London, Macmillan.

RECENTLY PUBLISHED "TEMI" (*)

No. 380 — *Bank Competition and Regulatory Reform: The Case of the Italian Banking Industry,* by P. ANGELINI and N. CETORELLI (October 2000).

No. 381 — *The Determinants of Cross-Border Bank Shareholdings: an Analysis with Bank-Level Data from OECD Countries,* by D. FOCARELLI and A. F. POZZOLO (October 2000).

No. 382 — *Endogenous Growth with Intertemporally Dependent Preferences*, by G. FERRAGUTO and P. PAGANO (October 2000).

No. 383 — *(Fractional) Beta Convergence*, by C. MICHELACCI and P. ZAFFARONI (October 2000).

No. 384 — *Will a Common European Monetary Policy Have Asymmetric Effects?*, by L. GUISO, A. K. KASHYAP, F. PANETTA and D. TERLIZZESE (October 2000).

No. 385 — *Testing for Stochastic Trends in Series with Structural Breaks*, by F. BUSETTI (October 2000).

No. 386 — *Revisiting the Case for a Populist Central Banke*r, by F. LIPPI (October 2000).

No. 387 — *The multimarket contacts theory: an application to italian banks,* by R. DE BONIS and A. FERRANDO (December 2000).

No. 388 — *La "credit view" in economia aperta: un'applicazione al caso italiano,* by P. CHIADES and L. GAMBACORTA (December 2000).

No. 389 — *The monetary trasmission mechanism: evidence from the industries of five OECD countries,* by L. DEDOLA and F. LIPPI (December 2000).

No. 390 — *Disuguaglianza dei redditi individuali e ruolo della famiglia in Italia,* by G. D'ALESSIO and L. F. SIGNORINI (December 2000).

No. 391 — *Expectations and information in second generation currency crises models,* by M. SBRACIA and A. ZAGHINI (December 2000).

No. 392 — *Unobserved Factor Utilization, Technology Shocks and Business Cycles,* by D. J. MARCHETTI and F. NUCCI (February 2001).

No. 393 — *The Stability of the Relation between the Stock Market and Macroeconomic Forces,* by F. PANETTA (February 2001).

No. 394 — *Firm Size Distribution and Growth,* by P. PAGANO and F. SCHIVARDI (February 2001).

No. 395 — *Macroeconomic Forecasting: Debunking a Few Old Wives' Tales,* by S. SIVIERO and D. TERLIZZESE (February 2001).

No. 396 — *Recovering the Probability Density Function of Asset Prices Using GARCH as Diffusion Approximations,* by F. FORNARI and A. MELE (February 2001).

No. 397 — *A Simple Approach to the Estimation of Continuous Time CEV Stochastic Volatility Models of the Short-Term Rate,* by F. FORNARI and A. MELE (February 2001).

No. 398 — *La convergenza dei salari manifatturieri in Europa,* by P. CIPOLLONE (February 2001).

No. 399 — *Labor Income and Risky Assets under Market Incompleteness: Evidence from Italian Data*, by G. GRANDE and L. VENTURA (March 2001).

No. 400 — *Is the Italian Labour Market Segmented?*, by P. CIPOLLONE (March 2001).

No. 401 — *Optimal Debt Maturity under EMU,* by R. GIORDANO (March 2001).

No. 402 — *Il modello di specializzazione internazionale dell'area dell'euro e dei principali paesi europei: omogeneità e convergenza,* by M. BUGANELLI (March 2001).

No. 403 — *Seasonality and Capacity: an Application to Italy,* by G. DE BLASIO and F. MINI (June 2001).

No. 404 — *The Effects of Bank Consolidation and Market Entry on Small Business Lending,* by E. BONACCORSI DI PATTI and G. GOBBI (June 2001).

(*) Requests for copies should be sent to:
Banca d'Italia – Servizio Studi – Divisione Biblioteca e pubblicazioni – Via Nazionale, 91 – 00184 Rome
(fax 0039 06 47922059). They are available on the Internet at www.bancaditalia.it