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THE DEMAND FOR M3 IN THE EURO AREA

G. COENEN* AND J.-L. VEGA
Directorate General Research, European Central Bank, Frankfurt am Main, Germany

SUMMARY

In this paper, an empirically stable money demand model for M3 in the euro area is constructed. Starting with a multivariate system, three cointegrating relationships with economic content are found: (i) the spread between the long-term and the short-term nominal interest rates, (ii) the long-term real interest rate, and (iii) a long-run demand for broad money M3. There is evidence that the determinants of M3 money demand are weakly exogenous with respect to the long-run parameters. Hence, following a general-to-specific modelling approach, a parsimonious conditional error-correction model for M3 money demand is derived which can be interpreted economically. For the conditional model, long-run and short-run parameter stability is extensively tested and not rejected. Copyright © 2001 John Wiley & Sons, Ltd.

1. INTRODUCTION

In October and December 1998, the Governing Council of the European Central Bank announced the key elements of its monetary policy strategy. As explained in ECB (1999a), these comprise a quantitative definition of the primary objective, namely price stability, and the 'two pillars' used for achieving this objective: a prominent role for money, as signalled by the publication of a reference value for the growth rate of broad money M3, and a broadly based assessment of the outlook for, and risks to, price stability in the euro area. Issing *et al.* (2001) discuss the analytical foundations of the monetary policy strategy of the ECB.

While the role assigned to money in the strategy is primarily based on theoretical grounds (namely the claim that inflation is ultimately a monetary phenomenon), empirical evidence, together with conceptual considerations, may indeed play an important role in selecting the particular monetary aggregate that best serves the purposes at hand. The existence of a stable and predictable relationship between the demand for a given monetary aggregate and its macroeconomic determinants together with its leading indicator and controllability properties have traditionally been considered—as noted in ECB (1999b)—key elements in this respect. Besides these considerations, there are plenty of good reasons for paying attention to the empirical properties of money demand. As Goldfeld (1994) argues: 'The relation between the demand for money balances and its determinants is a fundamental building block in most theories of macroeconomic behaviour and is a critical component in the formulation of monetary policy.'

Against this background, this paper presents some results of recent empirical research carried out at the European Central Bank on the demand for broad money M3 in the euro area. The paper is organized as follows. Section 2 briefly discusses our benchmark long-run specification for the estimation of the demand for broad money M3 in the euro area and the data underlying

* Correspondence to: Günter Coenen, European Central Bank, Kaiserstr. 29, D-60311 Frankfurtam Main, Germany. E-mail: gunter.coenen@ecb.int

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the empirical analysis. Section 3 investigates the cointegration properties of the data by means of the application of the Johansen procedure to a set of variables consisting of real holdings of M3 ($m - p$), real GDP (y), short-term (RS) and long-term (RL) nominal interest rates and the inflation rate as measured by the annualized quarterly change in the log of the GDP deflator ($\pi/4 = \Delta p$). In the light of the empirical findings, Section 4 develops a conditional model for M3 money demand in the euro area. We proceed in two steps and follow a general-to-specific modelling approach. First, an unrestricted autoregressive distributed lag (ADL) model in $m - p$, y, RS, RL and π is estimated and its long-run solution computed; second, the results obtained in the first step are then used for deriving a parsimonious, economically interpretable, conditional error-correction model for $\Delta(m - p)$.[1] Finally, Section 5 draws the conclusions from the analysis.

2. THE ECONOMIC MODEL AND THE DATA

2.1. The Model

While money is held for a number of purposes,[2] most theories of money demand lead, as argued in Ericsson (1999), to a long-run specification of the form:

$$M^d/P = g(Y, \tilde{R}) \tag{1}$$

where M^d, P, Y and $\tilde{R}$ stand for nominal money, the price level, a scale variable and a vector of returns on various assets. In applied work, a (semi-) log-linear form is often found to be an acceptable empirical approximation to equation (1), namely:

$$m_t^d - p_t = \gamma_0 + \gamma_1 y_t + \gamma_2 R_t^{\text{own}} + \gamma_3 R_t^{\text{out}} + \gamma_4 \pi_t \tag{2}$$

where variables in lower case indicate logs, $\pi/4 = \Delta p$, and R^{own} and R^{out} stand, respectively, for the nominal rates of return on financial assets included in and excluded from the definition of the monetary aggregate.

In equation (2) above, γ_1 measures the long-run elasticity of money demand with respect to the scale variable, while γ_2, γ_3 and γ_4 are, in turn, the long-run semi-elasticities with respect to the own and alternative rates of return and with respect to the inflation rate. Expected signs for the parameters in equation (2) are $\gamma_1 > 0$, $\gamma_2 > 0$, $\gamma_3 < 0$, $\gamma_4 < 0$ and, possibly, $\gamma_2 = -\gamma_3$. In the latter case, long-run money demand can be expressed as a function of the spread $R^{\text{out}} - R^{\text{own}}$, which in turn is interpretable as the opportunity cost of holding money.

Long-run price homogeneity of money demand has been assumed in (2), as predicted by most theories, but this can be empirically tested. Some theories also predict particular values for γ_1. For instance, $\gamma_1 = 0.5$ in the Baumol–Tobin model or $\gamma_1 = 1.0$ under some formulations of the quantity theory of money. Values $\gamma_1 > 1.0$ are also found with some frequency in the empirical literature for broad definitions of money, which in turn is customarily interpreted as proxying omitted wealth effects in equation (2). Extension of equation (2) to include wealth can be justified

[1] All computations throughout the paper are made using PcGive 9.30.
[2] Traditionally, a number of distinct motives for holding money are pointed out in the literature, giving rise to a transactions demand, a precautionary demand and a speculative demand for money. See Goldfeld and Sichel (1990) and Laidler (1993).

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under a standard portfolio approach to asset demand theory. This is not pursued here, however, due to the lack of reliable wealth data for the euro area.

The inclusion of the inflation rate in equation (2) is the subject of some ongoing controversy in the literature. However, separate consideration of the inflation rate in dynamic models of money demand may be of particular interest for a number of reasons. First, it permits a reparameterization of the models in terms of real money holdings and the inflation rate. Such reparameterization allows for the theoretically plausible hypothesis of long-run price homogeneity of money demand but does not impose any untested (and frequently empirically rejected) common factor restriction of short-run price homogeneity. In the context of cointegrated systems, it may also permit some convenient simplifications when the money stock and the price level are found to be CI(2,1), i.e. m and p are I(2) but $m - p$ is I(1), such that the I(2) system can be mapped—as in Johansen (1992)—into an I(1) system.

Second, numerous authors have forcefully argued for the inclusion of the inflation rate as an important determinant of constant-parameter empirical models of money demand.[3] This is customarily justified—as argued in Ericsson (1999)—on the basis that it represents the opportunity cost of holding money rather than real assets. On different grounds, within a cost-minimizing framework similar to that in Hendry and von Ungern-Sternberg (1981), Wolters and Lütkepohl (1997) show that, in the presence of a short-run nominal adjustment mechanism and under inflation persistence, the inflation rate may enter the long-run relation even if it does not appear in the desired long-run demand for money function.

Third, it could be argued that the inclusion or exclusion of inflation in models of real money demand is an issue of dynamic specification to be settled at the empirical level. In this sense, while some ambiguity would necessarily remain on the interpretation of the role of the inflation rate, the consideration of inflation as one of the variables entering the long-run demand for money or, alternatively, affecting only the process of dynamic adjustment to the long-run equilibrium would have little empirical content, since—as argued in Goldfeld and Sichel (1987)—both interpretations lead to observationally equivalent empirical models.

2.2. The Data

The long-run specification given by equation (2) is our maintained hypothesis for estimation of the demand for broad money M3 in the euro area. Following earlier work by Fagan and Henry (1999), the following empirical counterparts proxying the variables in the r.h.s. of equation (2) were chosen: real GDP (y) for the scale variable, the GDP deflator (p) for the price level, the three-month money market rate (RS) for the return on assets included in the definition of M3, and the 10-year government bond yield (RL) for the return on assets excluded from the monetary aggregate. The choice of real GDP and the GDP deflator as the scale and price variables in the money demand function is standard in existing empirical work, though alternative measures such as total final expenditure, consumption or wealth are also frequently found. The choice of the short-term and long-term interest rates can be justified on the basis of the broadness of the M3 aggregate, which includes assets that are remunerated at or close to market rates, though alternative

[3] For instance, seven out of the fourteen articles published in Lütkepohl and Wolters (1999) include inflation as a determinant of the long-run demand for money.

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measures of the own rate deserve being explored as longer area-wide time series on deposit rates become available.[4]

The data on y, p, RS and RL are taken from the ECB area-wide model database. Their construction—details of which can be found in Fagan, Henry and Mestre (2001)—can be briefly summarized as follows. Euro area real GDP and the GDP deflator are seasonally adjusted and obtained from Eurostat for the period in which these time series are available; for earlier periods, they are calculated from different national sources and are then aggregated using fixed weights based on 1995 GDP at PPP rates. The short-term and long-term interest rates are obtained from the BIS database as weighted averages of national rates using also fixed weights based on 1995 GDP at PPP rates.

As defined in December 1998 by the Governing Council of the ECB, M3 consists of holdings by euro area residents of currency in circulation plus certain liabilities issued by Monetary Financial Institutions (MFIs) and, in the case of deposits, liabilities issued by some institutions which are part of the central government (such as post offices and treasuries). These include: overnight deposits, deposits with agreed maturity up to two years, deposits redeemable at notice up to three months, repurchase agreements, money market fund shares, money market paper, and debt securities with maturity up to two years. A detailed description of euro area monetary aggregates can be found in ECB (1999b).

As from January 1999, the ECB has regularly been publishing data on euro area monetary aggregates (M1, M2 and M3) denominated in euro. The monetary aggregates are compiled on the basis of the consolidated balance sheet of the MFI sector from data collected under the new harmonized system of money and banking statistics. This consolidated balance sheet in euro is, in turn, available with the same degree of detail only back to September 1997. For periods prior to September 1997, the ECB is not in a position to produce historical data on euro area monetary aggregates according to its regular compilation procedures and, therefore, longer time series can only be constructed on the basis of the aggregation of estimated national contributions to the euro area aggregates compiled from a number of not fully harmonized national statistical sources,[5] including—as far as information is available—cross-border positions of MFIs within the euro area. In the February 1999 issue of the Monthly Bulletin the ECB released some historical estimates back to 1980. The time series were expressed in euro, with the estimated historical national contributions aggregated being converted into the single currency using the irrevocable conversion rates fixed on 31 December 1998.

It is fair to say that there is no uncontroversial aggregation method for linking euro area pre-1999 and post-1999 data, reflecting the fundamental problem that it is only from 1999 onwards that a single currency is in place. The existing empirical literature on area-wide money demand has indeed addressed this issue and a number of methods have been employed in this respect. In Monticelli and Strauss-Kahn (1992) area-wide aggregates are converted into a common currency using current exchange rates vis-à-vis the ECU. In Fase and Winder (1999) fixed exchange rates

[4] Based on preliminary analysis, we did not find compelling evidence of the need to adjust RL for risk as in Baba, Hendry and Starr (1992). Therefore, we do not pursue this avenue below. We believe, however, that this issue may deserve being revisited in conjunction with the measurement of the own rate of M3. This will be a subject of future research.

[5] National contributions to euro area monetary aggregates need to be distinguished from old national monetary aggregates. In particular, the euro area definition of M3 does not coincide (in terms of asset coverage and sector/instrument/maturity classification) with national definitions of broad money in place during Stage Two. These national contributions have been estimated on the basis of national data which are compatible with the new system of money and banking statistics. A description of the construction of historical estimates can be found in the statistical annex to ECB (1999b).

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are employed. Fagan and Henry (1999) propose to weigh national aggregates according to a fixed weighting scheme based on GDP at PPP rates. Fase and Winder (1997) discuss the effects of alternative aggregation methods. Winder (1997) and Beyer, Doornik and Hendry (2001) provide for a general discussion on aggregation issues underlying the construction of historical euro area data. The use of a consistent aggregation method for monetary aggregates, on the one hand, and for the r.h.s. variables in equation (2), on the other, is often stressed in the literature as an important requirement when developing empirical models of money demand in the euro area.

Against the background of the choice of r.h.s. variables described above, we proceed as follows throughout the paper. From September 1997 onwards, M3 as regularly published by the ECB is always employed. For the period prior to September 1997, the fixed-weights method proposed in Fagan and Henry (1999) is used to aggregate the historical national contributions and to produce back-estimates of area-wide M3 starting from September 1997 levels. This method of aggregation ensures full compliance with the conceptual consistency requirement highlighted above and, therefore, it is used throughout the paper, bar Section 4.3.[6] This way of proceeding, however, has one drawback, namely that it departs from the compilation procedures for M3 based on the use of fixed conversion rates which are in place since the start of Stage Three. As this may potentially have some implications if the estimated models were to be used out of the estimation sample, we explore in Section 4.3 the effects that the aggregation of the estimated national contributions by means of fixed conversion rates (as published in ECB, 1999b) would have on our estimates when no change is introduced in the way in which the r.h.s. variables in equation (2) are computed.

Finally, it needs to be born in mind that the use of euro area historical data for monetary aggregates compiled using a set of across-countries common specifications which are as compatible as possible with the sector/instrument/maturity classification laid down in the new system of euro area money and banking statistics, marks a significant departure from the data used in previous empirical research on area-wide money demand. Regardless of the aggregation issue, the new data set minimises the risks that the move to the new system of statistics introduces a sizeable break in the estimates, rendering empirical evidence useless in the new context. On the other hand, the new data set makes it difficult to compare the results with previous analyses and, accordingly, this avenue is not pursued in this paper.

Figures 1(a) and 1(b) plot the time series used in the empirical analysis:[7] real holdings of M3 and real GDP (see Figure 1(a)), the short-term and long-term nominal interest rates and the inflation rate measured by the annualized quarterly change in the GDP deflator (see Figure 1(b)). From visual inspection, the trending behaviour which often characterizes non-stationary series is apparent in the plots of $m - p$ and y. Furthermore, RS, RL and π appear to share a common, possibly stochastic, nominal trend during the sample under investigation. The latter observation is more visible in Figures 1(c)–1(e), which depict time series for the long-term real interest rate (see Figure 1(c)), the short-term real rate (see Figure 1(d)) and the spread between the long-term and the short-term nominal rates (see Figure 1(e)). All of them look much more stationary than their individual counterparts in Figure 1(b) and the application

[6] Alternatively, historical series for euro area nominal and real GDP at fixed conversion rates could be computed and the corresponding implicit GDP deflator be calculated. While ensuring consistent aggregation, this procedure cannot be applied to interest rates. Moreover, area-wide GDP inflation so defined would, in general, be different from the weighted average of member countries' inflation rates, which is the concept most used in economic analysis.

[7] The lack of reliable non-seasonally adjusted series for area-wide real GDP and the GDP deflator constrains our choice between seasonally adjusted and unadjusted data. Thus, quarterly averages of seasonally adjusted M3 monthly data are used in the analysis. Seasonal adjustment has been made using SEATS.

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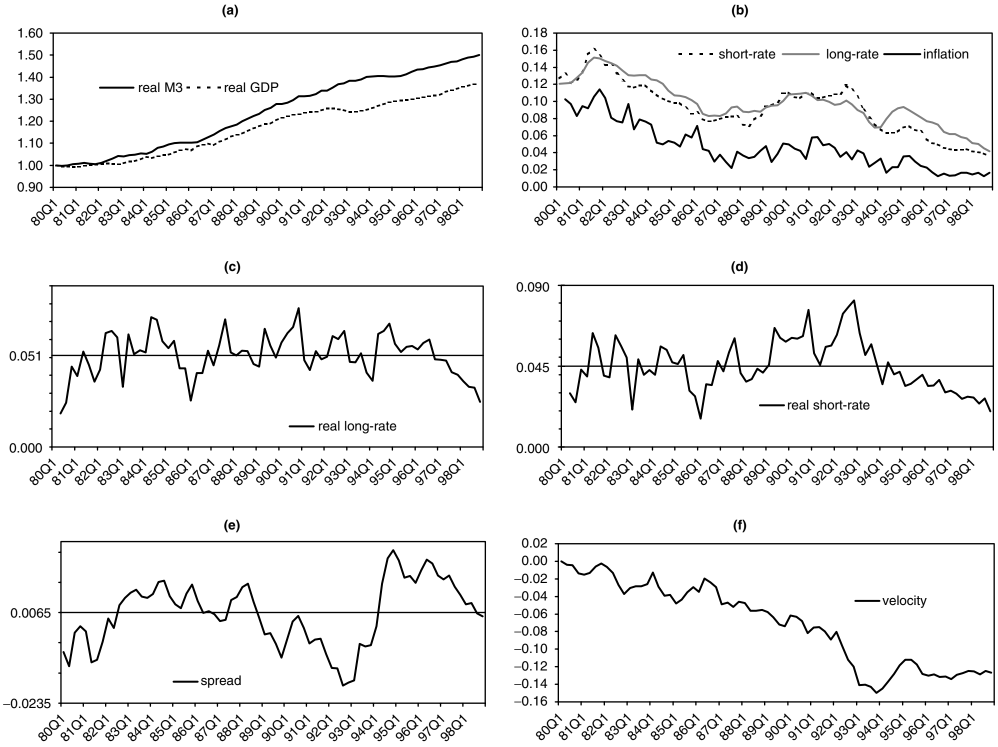

Figure 1. The data

of the Johansen procedure below will confirm that they indeed constitute simple cointegrating relationships among our set of variables. Finally, Figure 1(d) plots the income velocity of M3, showing the downward trending behaviour which has been reported elsewhere. Income velocity of M3 has declined by a cumulated 13% since the early 1980s, representing a -0.7% annual decline per year. It appears to have stabilized, however, in the most recent years. This pattern in velocity parallels to some extent the developments in area-wide inflation shown in Figure 1(b).

ADF tests for unit roots support the view that $m-p$, y, RS, and RL are I(1) for the sample under investigation. The same outcome applies to income velocity $y+p-m$. Less clear-cut results are obtained, however, for m and p. On the one hand, ADF tests tend to reject a second unit root for the full sample under consideration when the alternative contains a deterministic trend, i.e. Δm and Δp could be trend-stationary. On the other hand, recursive computation of the corresponding ADF tests indicates that this finding is not robust to the choice of sample and provides support for assuming that both variables can be better described as I(2). Additionally, since univariate tests are known to have low power against some stationary alternatives, multivariate tests for stationarity will be conducted as well within the application of the Johansen procedure in Section 3. The results therein provide further empirical support for the hypothesis that the inflation rate is not stationary during the period analyzed.

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3. THE COINTEGRATION ANALYSIS

In this section the cointegration properties of the data are investigated by means of the Johansen procedure (see Johansen, 1995) to the set of variables $z = (m - p, y, RS, RL, \pi)'$. An unrestricted constant, allowing for a linear trend in the variables but not in the cointegrating relationships and a dummy variable (*DUM86*)[8] are also included in the system. The estimation sample spans the period 1980 : Q4 to 1998 : Q4. The monetary aggregate, the GDP deflator and real GDP are seasonally adjusted. The interest rates are measured as percent per annum, expressed as fractions.

The results from the application of the Johansen procedure are summarized in Table I. The top panel of the table reports the Akaike (*AIC*) and Hannan-Quinn (*HQ*) information criteria for the selection of lag-length (k) and various diagnostic statistics on the system residuals: Lagrange-multiplier tests for first-order (*LM(1)*), fourth-order (*LM(4)*) and up to eighth-order (*LM(1,8)*) residual autocorrelation, the Doornik and Hansen multivariate test for normality (*NORM*), and a White-type test for heteroskedasticity (*HET*). A description of the tests plus appropriate references can be found in Doornik and Hendry (2000).

According to the test statistics above, the VAR with $k = 2$ appears reasonably well specified over the estimation sample, although some residual non-normality is revealed by the Doornik and Hansen test. Excess kurtosis in the residuals of the equations for the short-term rate and, to a lesser extent, for real GDP, due to the presence of some outliers at the beginning of the estimation period, appears primarily responsible for this finding.

As regards cointegration, Table I shows the results for the trace test, since—as argued in Cheung and Lai (1993)—this is considered to be robust to the non-normality encountered in the data. Furthermore, as critical values are affected by the inclusion of dummy variables, the rank test statistics reported in Table I refer to the system both with and without *DUM86* (the latter in square brackets). Taking small sample corrections into account as suggested in Cheung and Lai (1993), the results point to the existence of three cointegrating vectors at the 90% confidence level. Conditional on the choice of cointegration rank $r = 3$, tests for long-run exclusion and stationarity of each variable as well as tests for structural hypotheses on β are reported in the table. The tests do not allow to reject at standard confidence levels the stationarity of both the spread between the long-term and the short-term interest rate (H_0^1), consistent with the expectations theory of the term structure of interest rates, and the real long-term interest rate (H_0^2), consistent with the Fisher parity. The stationarity of the real short-term interest rate is in turn implied by the non-rejection of the joint hypothesis $\mathrm{H}_0^1 \cap \mathrm{H}_0^2$. On the contrary, long-run homogeneity of real money and real income is rejected at standard confidence levels, both when the hypothesis is tested in isolation and when it is tested jointly with $\mathrm{H}_0^1 \cap \mathrm{H}_0^2$. The tests yield $\chi_3^2 = 10.31(.016)$ and $\chi_5^2 = 12.07(.034)$ respectively (p-values in parentheses).

The estimated cointegrating vectors are reported in Table I in the form of irreducible cointegrating relations (IC), following Davidson (1998).[9] Besides the two over-identified relationships, namely the spread $(RL - RS)$ $[\beta_2' = (0, 0, -1, 1, 0)]$ and the real long-term rate $(RL - \pi)$

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[8] *DUM86* equals: 0.5 in the first, third and fourth quarters of 1986, 1 in the second quarter of 1986, and 0 elsewhere. This dummy variable is needed to account for special developments in German data around the period in which debt securities were subjected to reserve requirements. This led to substantial differences between the developments in the German contribution to area-wide M3 (which includes debt securities) and the historical German M3 aggregate (which does not include debt securities).

[9] A set of I(1) variables is called irreducibly cointegrated (IC) if they are cointegrated, but dropping any of the variables leaves a set that is not cointegrated.

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Table I. Summary of results from the application of the Johansen procedure: M3 (fixed GDP PPP rates). 1980 : Q4–1998 : Q4

Informat. criteria	$k = 1$	$k = 2$	$k = 3$	$k = 4$		
AIC	−53.29	−54.06	−53.65	−53.60		
HQ	−52.85	−53.30	−52.58	−52.20		
Test diagnostics	*LM(1)*	*LM(4)*	*LM(1,8)*	*N*	*HET*	
($k = 2$)	0.60	0.79	1.07	25.50**	0.86	
Cointegration rank	$r = 0$	$r <= 1$	$r <= 2$	$r <= 3$	$r <= 4$	
trace test	115.20	67.48	39.21	15.88	0.17	
	[101.9**]	[61.13*]	[32.2+]	[11.48]	[0.02]	
Chi-square tests for	$m - p$	y	RS	RL	π	
exclusion	26.95**	23.84**	16.95**	9.21*	22.70**	
stationarity	21.50**	22.04**	14.82**	19.11**	16.15**	
Tests for structural hypotheses	H_0^1	H_0^2	$H_0^1 \cap H_0^2$			
	(0,0,−1,1,0)	(0,0,0,1,−1)				
	2.47	3.01	7.46			
	(0.29)	(0.22)	(0.11)			
The cointegrating vectors	β_1	s.e.	β_2	s.e.	β_3	s.e.
$(m - p)_t$	1.000	—	—	—	—	—
y_t	−1.163	0.028	—	—	—	—
RS_t	1.291	0.117	−1.000	—	—	—
RL_t	—	—	1.000	—	1.000	—
π_t	—	—	—	—	−1.000	—
The loadings	α_1	t-ratio	α_2	t-ratio	α_3	t-ratio
$(m - p)_t$	−0.127	4.65**	−0.261	3.88**	0.098	1.68+
y_t	0.028	0.73	0.246	2.66**	−0.141	1.75+
RS_t	−0.060	1.50	0.020	0.21	−0.115	1.35
RL_t	−0.011	0.40	−0.048	0.69	−0.146	2.42*
π_t	−0.012	0.17	0.009	0.05	0.355	2.27*
	H_0^3		$H_0^1 \cap H_0^2 \cap H_0^3$			
	$(\alpha_{21} = \alpha_{31} = \alpha_{41} = \alpha_{51} = 0)$					
	2.99		10.44			
	(0.22)		(0.24)			

Note: the superscripts +, * and ** indicate rejections at 10%, 5% and 1% significance levels, respectively.

$[\beta_3' = (0, 0, 0, 1, -1)]$, an additional just-identified cointegrating vector $[\beta_1' = (1, -\delta, -\eta, 0, 0)]$ expressing real holdings of M3 as a function of the real and nominal stochastic trends driving the system is reported in the table:

$$(m - p)_t = const + \underset{(0.028)}{1.163}\, y_t - \underset{(0.117)}{1.291}\, RS_t + \hat{u}_t \tag{3}$$

(standard errors in parentheses).

Just-identification of β_1' has been obtained by setting arbitrarily to zero the coefficients of the long-term interest rate and the inflation rate. It should be noticed that, while this normalisation has been chosen as to make Equation (3) look like a traditional textbook money demand function, nothing so far guarantees its structural interpretation.[10] In particular, any linear combination

[10] As opposed to a solved form, following the terminology in Davidson (1998).

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$\beta = \sigma_1\beta_1 + \sigma_2\beta_2 + \sigma_3\beta_3$ is also a cointegrating relationship and, for some parameter values, would be a plausible candidate for constituting our relationship of interest. Normalizing in real money holdings ($\sigma_1 = 1$), $\beta' z_t$ can be written as follows:

$$(m - p)_t = k + \delta y_t + \eta RS_t - \sigma_2(RL_t - RS_t) - \sigma_3(RL_t - \pi_t) \tag{4}$$

$$= \gamma_0 + \gamma_1 y_t + \gamma_2 RS_t + \gamma_3 RL_t + \gamma_4 \pi_t \tag{5}$$

where economic theory suggests that money demand would correspond to Equation (5) with: $\gamma_1 > 0$; $\gamma_2 > 0$ (if the short-term rate proxies the own rate of money), $\gamma_3 < 0$ (if the long-term rate measures the return of financial assets alternative to those included in the monetary aggregate) and, possibly, $\gamma_2 = -\gamma_3$ (if the spread is the relevant opportunity cost of holding money relative to financial assets not included in the definition of money), and $\gamma_4 < 0$ (if money is a substitute for real assets). Equation (5) corresponds to our benchmark long-run specification for the estimation of the demand for broad money M3 in the euro area and the parameters in (5) constitute our long-run parameters of interest.

It is also straightforward to show that the parameters in equations (4) and (5) are related as follows:

$$\gamma_1 = \delta \tag{6}$$

$$\gamma_2 = \eta + \sigma_2 \tag{7}$$

$$\gamma_3 = -(\sigma_2 + \sigma_3) \tag{8}$$

$$\gamma_4 = \sigma_3 \tag{9}$$

From Equations (6) to (9) some results of interest concerning the parameters δ and η in the first cointegrating vector β_1 follow. First, the parameter δ identifies unambiguously the long-run income elasticity of money demand. Second, η identifies unambiguously the effect that the common nominal trend has on M3 real balances, i.e. $\eta = \gamma_2 + \gamma_3 + \gamma_4$. It does not identify the parameter measuring the semi-elasticity of money demand with respect to the short-term rate (γ_2) unless $\gamma_3 + \gamma_4 = 0$, which runs against economic intuition since both γ_3 and γ_4 are expected to be of negative sign. Third, if the two interest rates entered the long-run money demand as a spread ($\gamma_2 = -\gamma_3$), η would correspond to the long-run semi-elasticity of money demand with respect to inflation ($\eta = \gamma_4$). And finally, if the inflation rate does not enter the long-run demand for money function ($\gamma_4 = 0$), the results reported in Table I (i.e. the significance of η) would rule out the spread formulation of long-run money demand.

Interestingly, weak exogeneity of real GDP, the short-term and long-term interest rates and the inflation rate when the parameters of interest are those of the first cointegrating vectors is not rejected by the tests reported in Table I (H_0^3).[11] One implication of this result is that, given the structure of the estimated cointegration space, weak exogeneity is not rejected either for the parameters of any linear combination involving β_1, on the one hand, and the two other over-identified cointegrating vectors β_2 and β_3, on the other. This holds in particular for the parameters of the long-run demand for money function (5) which, as shown above, can be expressed as such a linear combination. That in turn indicates that, as far as the parameters of the long-run money demand are concerned, nothing can be learnt from the equations in the system other than

[11] Estimation of δ and η under $H_0^1 \cap H_0^2 \cap H_0^3$ provides the following results: $\delta = 1.134$ (0.029) and $\eta = -1.475$ (0.124).

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the equation for real balances. Accordingly, efficient inference on the parameters of long-run money demand will be made in the next section on the basis of a parsimonious, conditional, single-equation model for area-wide money demand.[12]

4. A DYNAMIC MODEL OF MONEY DEMAND IN THE EURO AREA

In view of the results on cointegration and weak exogeneity obtained from the application of the Johansen procedure in Section 3, a conditional model for M3 money demand in the euro area is developed in this section. We proceed in two steps and follow a general-to-specific modelling approach (see, for instance, Hendry, 1995). First, given the choice of lag length in the VAR, a second-order unrestricted autoregressive distributed lag (ADL) model in $m - p$, y, RS, RL and π is estimated and its long-run solution computed; and, second, the results obtained in the first step are then used for deriving a parsimonious, economically interpretable, conditional error-correction model for $\Delta(m - p)$.

The estimation period spans the sample 1980 : Q4 to 1998 : Q4 and particular attention is paid to recursive estimates over the most recent period. Since it has been argued that the introduction of the euro marks a significant regime shift which could affect empirical relationships estimated on the basis of past data (see, for instance, ECB, 1999a), analysis of money demand stability during the period immediately preceding the introduction of the single monetary policy should provide evidence on the empirical relevance of such risks by detecting any anticipated effect.

4.1. Long-run Money Demand

The long-run solution of the estimated ADL(2) model is given by:

$$(m - p)_t = const + \underset{(0.058)}{1.125}\, y_t - \underset{(0.359)}{0.865}\,(RL_t - RS_t) - \underset{(0.329)}{1.512}\,\pi_t + \hat{u}_t \tag{10}$$

where the spread restriction turned out to be statistically acceptable [$F(1, 57) = 1.66$ (0.20)], while the exclusion of the inflation rate and the unit income elasticity were rejected [$F(1, 58) = 21.18$ (0.00) and $F(1, 58) = 4.68$ (0.035) respectively]. Rejection of the latter is consistent with the findings in Section 3.

Equation (10) closely resembles our benchmark long-run specification and, following Boswijk (1995a), it is a natural candidate for structural interpretation as a money demand cointegrating relation.[13] It shows a simple specification for the long-run demand for M3 whereby holdings of M3 real balances are determined by a measure of the volume of transactions (as proxied by real income), the opportunity cost of holding money relative to financial assets not included in the definition of the monetary aggregate (as proxied by the spread between the long-term and short-term interest rates) and the inflation rate (proxied by the change in the GDP deflator).

The signs and magnitudes of the estimated long-run coefficients also appear quite plausible on theoretical grounds. The long-run income elasticity is estimated significantly above one (though not

[12] The results from the tests for weak exogeneity, however, need to be interpreted with caution since the 'weak exogeneity' status of a given variable may change when the system is augmented with additional variables. By contrast, cointegration is a property which is invariant to expansions of the system.
[13] See also Bårdsen and Fisher (1995). A careful discussion of Boswijk (1995a) can be found in Ericsson (1995) and a reply in Boswijk (1995b).

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far above, i.e. 1.13), a finding which—as argued in Section 2—is often interpreted as proxying omitted wealth effects in the demand for money function. Some evidence at the MU level on the relevance of wealth in the long-run demand for money function can be found in Fase and Winder (1997, 1999). The estimated long-run semi-elasticities with respect to the spread and the inflation rate are, respectively, −0.87 and −1.51 (though in both cases less precisely estimated than the income coefficient).[14] Equation (10) provides an explanation for the downward trending behaviour of income velocity of M3 depicted in Figure 1(f) in terms of two main factors: first, the income elasticity of money demand higher than one; and, second, the significant fall in the inflation rate witnessed by the euro area over the sample period. The spread between the long-term and the short-term interest rates, which shows no particular trending behaviour when the entire sample is considered, does not appear to contribute significantly to this long-term pattern of velocity. However, the low speed of adjustment of this variable to its estimated long-run equilibrium indicates that its contribution cannot be disregarded at medium-term horizons.

The income elasticity and the inflation semi-elasticity do not differ significantly from the estimates obtained within the system approach in Section 3, providing further evidence on the validity of weak exogeneity when the parameters of interest are the long-run parameters of money demand. In this latter respect, formal statistical tests do not reject at standard confidence levels that the long-run relationship given by equation (10) lies in the cointegration space estimated within the system approach used in Section 3. The tests yield $\chi^2_2 = 1.52$ (0.47) (if the cointegration space is estimated unrestrictedly), $\chi^2_6 = 7.85$ (0.25) (if the cointegration space is estimated under $H_0^1 \cap H_0^2$) and $\chi^2_{10} = 10.48$ (0.40) (if the cointegration space is estimated under $H_0^1 \cap H_0^2 \cap H_0^3$).

The parameters of equation (10) above turn out to be pretty stable in recent times. In this connection, panels 9 to 12 in Figure 2 show the recursive estimates from 1993 : Q4 onwards of the long-run parameters of real GDP, the spread, inflation and the price level in (10) together with plus/minus twice their standard errors. As regards the latter, the recursive estimates suggest that long-run price homogeneity, which was assumed in Section 2 and which is maintained throughout the paper, is an acceptable characterization of our estimates of long-run money demand.

4.2. A Dynamic Model of Money Demand

While the unrestricted ADL model is valuable for computing the long-run equation (10), it is clearly over-parameterized for many other purposes. Therefore, we proceed in a second step to estimate a parsimonious dynamic model for money demand which is both economically interpretable and statistically acceptable. The estimation results for the period 1980 : Q4–1998 : Q4 are summarized in equation (11) below (t-ratios in parentheses):

$$\begin{aligned}\Delta(m-p)_t = {} & \underset{(10.99)}{-0.690} + \underset{(1.78)}{0.071}\,\Delta^2 y_t + \underset{(2.65)}{0.194}\,\Delta_2 RS_t/2 - \underset{(4.47)}{0.353}\,\Delta RL_{t-1} - \underset{(10.79)}{0.526}\,\Delta_2\pi_t/2 \\ & - \underset{(11.10)}{0.132}\,[(m-p) - 1.125y + 0.865(RL-RS) + 1.512\pi]_{t-2} \\ & - \underset{(5.00)}{0.0095}\,DUM86_t + \hat{v}_t \end{aligned} \tag{11}$$

[14] The relative size of these two coefficients is somewhat counter-intuitive if the latter were to be interpreted as the sensitivity of money demand to non-financial assets. As argued in Section 2.1, however, this is not the only possible interpretation for the inclusion of the inflation rate in dynamic models of money demand.

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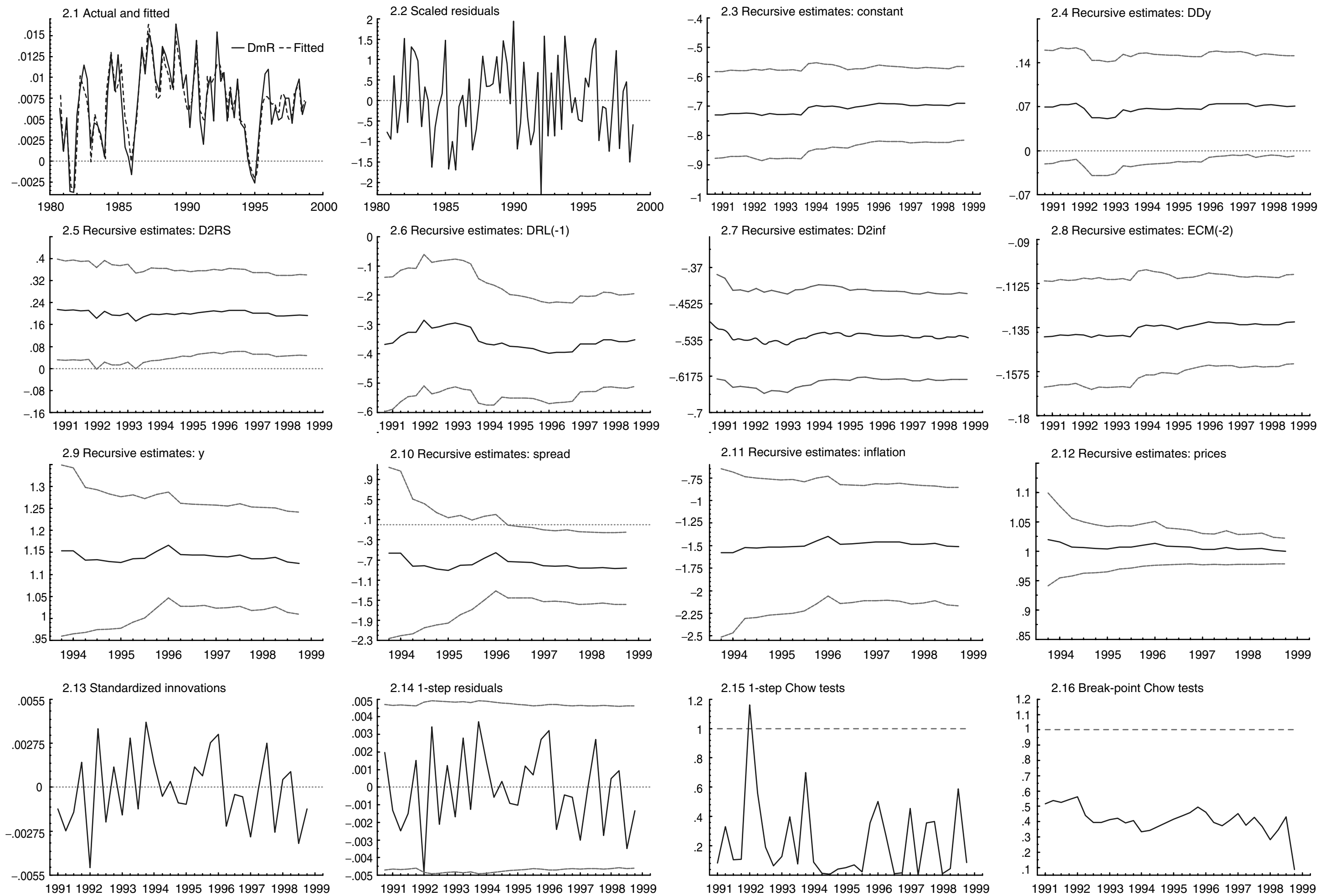

Figure 2. Graphical evaluation and recursive estimates: fixed GDP PPP weights

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$T = 73$ (1980 : Q4–1998:Q4)	$R^2 = 0.80$	$\sigma = 0.230\%$	$DW = 2.18$
$LM(1) = 0.649(0.42)$	$LM(4) = 0.319(0.57)$	$LM(1,4) = 0.734(0.57)$	
$ARCH(4) = 0.532$ (0.71)	$HET = 0.58$ (0.85)	$NORM = 1.53$ (0.47)	
$RESET = 0.158$ (0.69)	$RED = 0.44$ (0.82)	$HANS^1 = 0.041$	$HANS^2 = 0.622$

where: *LM(i)* and *LM(1,i)* stand for the Lagrange multiplier F-tests for residual auto-correlation of order i and up to the ith order respectively; *ARCH* is the Engle F-test for autoregressive conditional heteroskedasticity; *HET* is the White F-test for heteroskedasticity; *NORM* is the Doornik and Hansen χ^2-test for normality; *RESET* is the regression specification F-test due to Ramsey; *RED* tests whether model (11) parsimoniously encompasses the unrestricted ADL(2) model; and $HANS^1$ and $HANS^2$ are the Hansen tests for variance and parameter within-sample stability. P-values are reported in parentheses. A description of the tests plus appropriate references can be found in Hendry and Doornik (1999). Panels 1 and 2 in Figure 2 record the time series of fitted and actual values for $\Delta(m - p)$ and the scaled residuals from the model.

From a statistical point of view, the estimated model appears well specified, with tests showing no signs of residual autocorrelation, heteroscedasticity or non-normality. The reductions involved in moving from the unrestricted ADL(2) to the parsimonious model given by equation (11) are not rejected either (*RED*) and, therefore, equation (11) parsimoniously encompasses the ADL(2). The coefficient of the equilibrium correction term is highly significant, indicating that a long-run relationship exists between real holdings of M3, on the one hand, and real income, the spread and the inflation rate, on the other. The size of this coefficient, however, indicates that disequilibria are corrected only slowly. This is in line with empirical evidence for EU countries, as surveyed in Browne *et al*. (1997), and suggests that the costs of being out of the equilibrium (or alternatively the benefits of being in equilibrium) are small. The model parameters appear reasonably stable within the estimation sample, as indicated by the Hansen tests, and no major problems are detected when the equation is used for producing one-step-ahead forecasts over the last six quarters.

Following Urbain (1992), Wu–Haussman tests for weak exogeneity of real income (τ_1), the short-term rate (τ_2), the long-term rate (τ_3), and the inflation rate (τ_4) with respect to the short-run parameters of money demand were also conducted. That involves testing for the significance of the residuals from the marginal unrestricted reduced-form models for y, *RS*, *RL* and π in the money demand equation. None of the tests rejects at standard confidence levels the null of weak exogeneity, validating the single-equation approach adopted herein for estimation of money demand. Results are as follows: $\tau_1 = 0.231$ (0.632), $\tau_2 = 0.382$ (0.539), $\tau_3 = 0.138$ (0.711) and $\tau_4 = 0.512$ (0.477). The joint test yields: $\tau = 0.394$ (0.812).

As regards parameter stability, equation (11) was estimated recursively over the period 1990 : Q4–1998 : Q4. A graphical summary of the results is shown in Figure 2. The figure records the recursively estimated coefficients plus/minus twice their standard errors together with other relevant output from the recursive estimation. The latter includes: standardized innovations, one-step forecast errors plus/minus twice the recursively estimated equation standard error, one-step-ahead Chow tests and breakpoint Chow tests scaled by their 5% significance values. The estimated parameters appear constant and significant over most of the sample. The Chow tests do not reveal any major non-constancy either.[15]

[15] There are two large errors in 1992 : Q1 and 1992 : Q2 which are virtually equal in magnitude but opposite in sign. The same anomalies appear in the analysis in Section 4.3. While this may suggest a one-time mismeasurement in the level of

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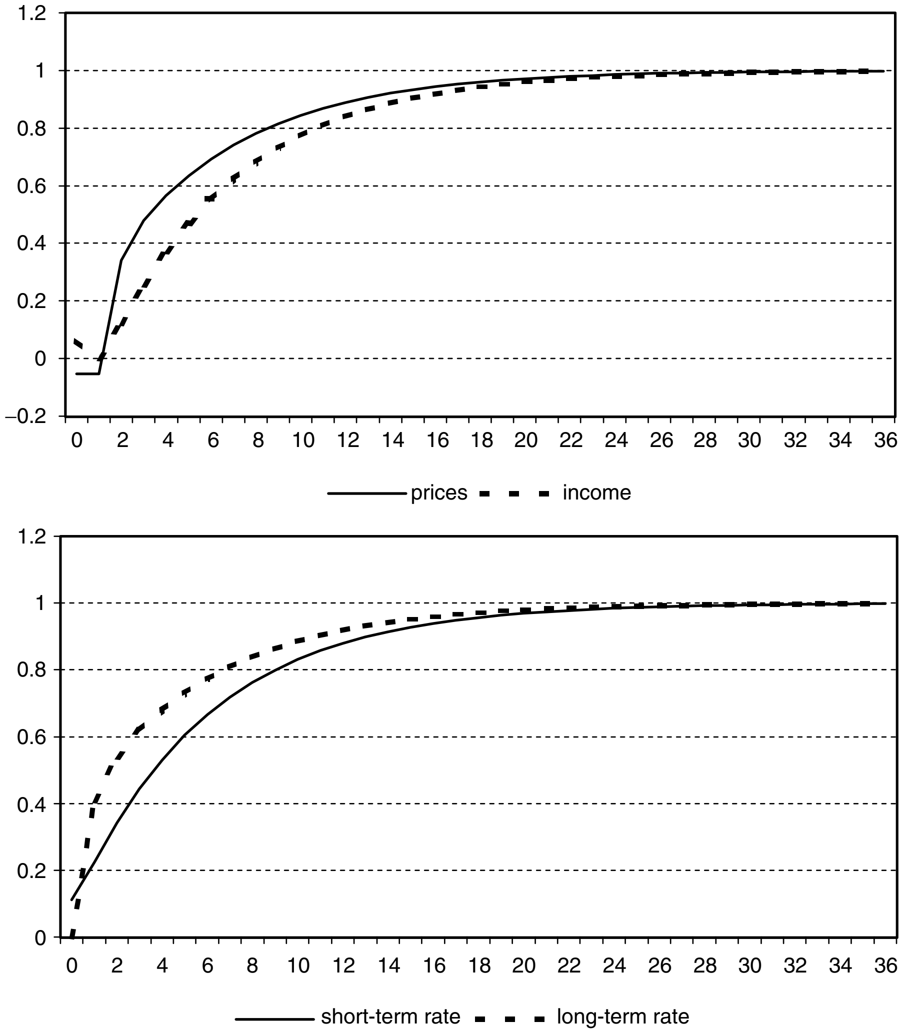

Figure 3. Cumulative normalized lag weights

Finally, with a view to gaining some insight into the dynamic properties of the estimated money demand equation, Figure 3 shows the cumulative normalised lag weights of prices, real income and the short-term and the long-term interest rates obtained from equation (11). The mean lags for p, y, RS and RL are 6.0, 7.6, 5.9 and 4.5 quarters respectively. The median lags are shorter: 3.2, 5.3, 3.6 and 1.7 quarters respectively. The short-run elasticities of nominal money with respect to prices are close to zero, in contrast with the estimated long-run elasticities. This result, which is often found in the empirical literature on money demand, is consistent—as argued in Ericsson

the money stock (or the price level) in 1992 : Q1, we prefer not to include any dummy due to the lack of corroborative explanation in this respect. Inclusion of a dummy does not appear to have any noticeable effect in the estimation of (10) or (11) other than the reduction of the standard errors of equation (11).

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(1999)—with Ss-models of money demand, with short-run factors determining the changes in money holdings given desired upper and lower bands and longer-run factors determining the bands themselves. Money demand is estimated to respond more rapidly to changes in the alternative rate than to changes in the own rate. That feature—as argued in Cabrero *et al.* (1998)—is of particular interest if the central bank were to exert some short- to medium-term control on the money stock. It should be born in mind, however, that the experiment conducted in computing the lag weights involves a rather unrealistic assumption, namely that the r.h.s. variables are orthogonal. Therefore, policy conclusions can only be drawn within more realistic multivariate settings which allow for the interplay of the different variables in the system.[16]

4.3. Some Results on Aggregation

As pointed out in Section 2.2, a number of methods for the calculation of historical aggregated series for the euro area have been employed in existing empirical work. It is also fair to say that any choice between them is somewhat arbitrary, reflecting the fundamental problem that it is only from 1999 onwards that a single currency is in place. In particular, the time series for area-wide M3 employed above is calculated backwards from September 1997 M3 levels in euro, using an output-weighted average of participating countries' developments. As discussed, this method of aggregation provides for a coherent historical calculation of euro area monetary aggregates, on the one hand, and the r.h.s. variables in the money demand equation, on the other. Of particular interest is that the resulting growth rates for M3 are consistent with the concept of area-wide inflation most widely used in economic analysis, i.e. a weighted average of countries' inflation rates.

One drawback, however, is that the aggregation method departs from the compilation procedures for euro area monetary aggregates which are in place since the start of Stage Three, whereby national data are converted into a common currency, the euro, using the irrevocable conversion rates announced on 31 December 1998. Since this may potentially have implications if the estimated models were to be used in the new context, we explore below the effect that the use of euro area M3 compiled on the basis of fixed conversion rates (as published in ECB, 1999b) may have on our estimates while noting that a discrepancy with respect to the way in which the remaining variables in our analysis are calculated is introduced.

On a conceptual basis, M3 growth figures resulting from both aggregation procedures can be interpreted as weighted averages of the estimated historical national contributions with weights depending on the method employed. Under the first aggregation method, weights are fixed and are calculated as countries' shares of area-wide GDP evaluated at PPP rates. Under the second, weights are time varying and depend on countries' shares on area-wide money M3 evaluated at fixed conversion rates. On average, the second method gives more weight to countries with higher liquidity ratios.

At the empirical level, a first conclusion that can be drawn from visual inspection of Figure 4 is that the difference between the two aggregation procedures should not be overstated. The figure plots the levels (see Figure 4(a)) and the quarterly growth rates (see Figure 4(b)) of M3 resulting from the application of the two aggregation methods. Both series in levels move closely together for most of the sample period with differences mainly concentrated at the beginning of the 1980s. The differences in the quarterly growth rates are of moderate size in statistical terms when compared,

[16] The interested reader is referred to Section 5 of the working paper version of the present article for an impulse-response analysis within the estimated VECM (http://www.ecb.int).

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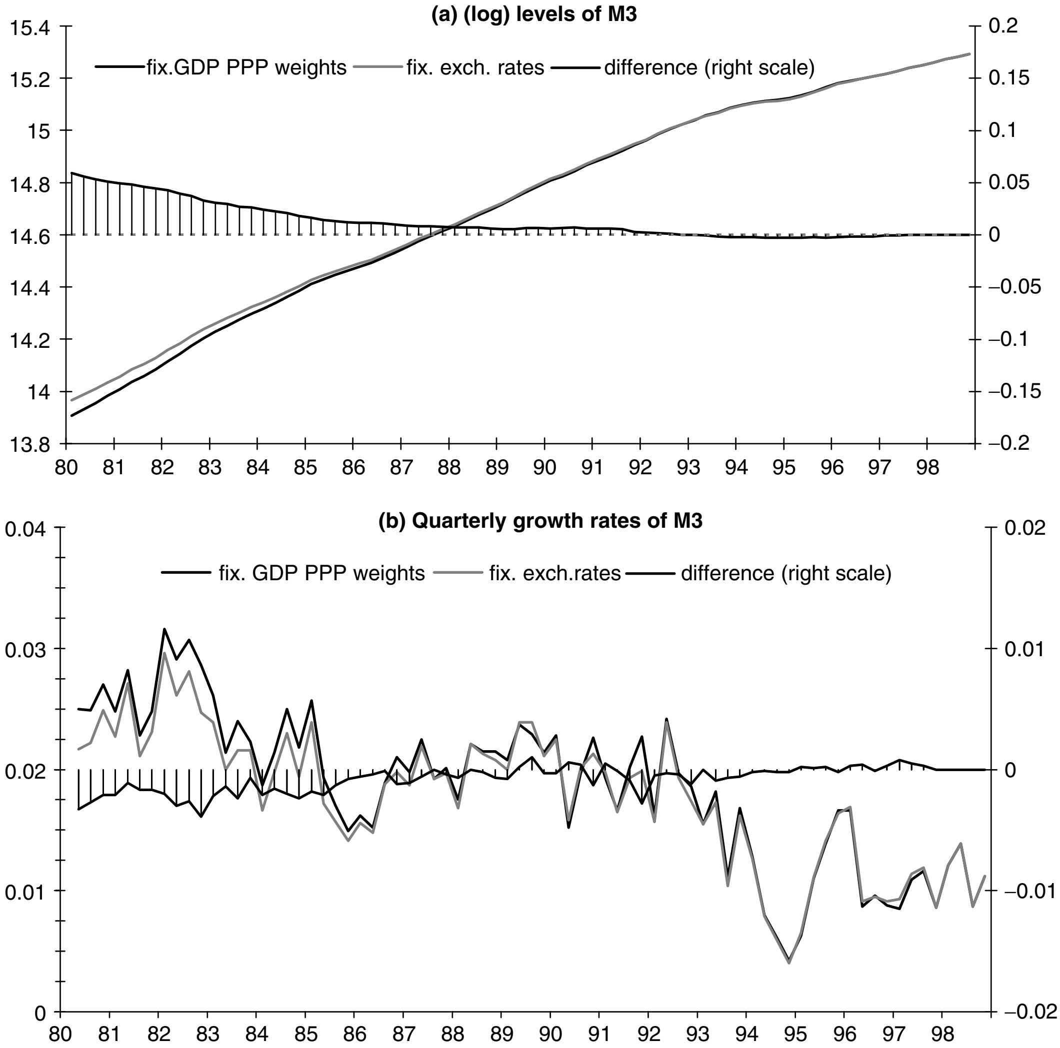

Figure 4. The monetary aggregates

for instance, with the standard error reported in equation (11) above. They are quite persistent, however, implying long-lasting effects when the levels of the series are considered. Thus, during the first half of the eighties the output-weighted M3 increased around 5% more in cumulative terms than the corresponding aggregate computed using the fixed conversion rates.

Formal analysis using the Johansen procedure confirms that both time series do cointegrate with cointegrating vector (1, −1). That in turn suggests that the cointegration results reported in Section 3 should be little affected by the change in the aggregation method for M3. Table II confirms this intuition by reproducing the analysis in Section 3 and Table I for the monetary aggregate M3 computed using the fixed conversion rates. None of the conclusions in Section 3 are changed by the evidence presented in the table.

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Table II. Summary of results from the application of the Johansen procedure: M3 (fixed conversion rates). 1980 : Q4–1998 : Q4

Informat. criteria	$k = 1$	$k = 2$	$k = 3$	$k = 4$		
AIC	−53.25	−54.02	−53.66	−53.62		
HQ	−52.80	−53.26	−52.58	−52.23		
Test diagnostics	*LM(1)*	*LM(4)*	*LM(1,8)*	*N*	*HET*	
($k = 2$)	0.68	0.88	1.08	25.53^{**}	0.81	
Cointegration rank	$r = 0$	$r <= 1$	$r <= 2$	$r <= 3$	$r <= 4$	
trace test	110.9	69.09	39.62	15.44	0.23	
	[100.6^{**}]	[63.65^{*}]	[31.12^{+}]	[11.37]	[0.03]	
Chi-square tests for	$m - p$	y	RS	RL	π	
exclusion	24.04^{**}	21.93^{**}	13.84^{**}	8.37^{*}	16.32^{**}	
stationarity	22.87^{**}	22.89^{**}	16.04^{**}	20.71^{**}	17.56^{**}	
Tests for structural hypotheses	H_0^1	H_0^2	$H_0^1 \cap H_0^2$			
	(0,0,−1,1,0)	(0,0,0,1,−1)				
	1.52	4.09	9.22^{+}			
	(0.47)	(0.13)	(0.06)			
The cointegrating vectors	β_1	s.e.	β_2	s.e.	β_3	s.e.
$(m - p)_t$	1.000	—	—	—	—	—
y_t	−1.156	0.030	—	—	—	—
RS_t	0.881	0.128	−1.000	—	—	—
RL_t	—	—	1.000	—	1.000	—
π_t	—	—	—	—	−1.000	—
The loadings	α_1	t-ratios	α_2	t-ratios	α_3	t-ratios
$(m - p)_t$	−0.134	5.25^{**}	−0.183	3.48^{**}	0.010	0.21
y_t	0.022	0.61	0.217	2.90^{**}	−0.121	1.76^{+}
RS_t	−0.044	1.13	0.071	0.89	−0.160	2.18^{*}
RL_t	−0.004	0.14	−0.032	0.56	−0.158	3.03^{**}
π_t	0.047	0.66	0.091	0.62	0.314	2.34^{*}
	H_0^3		$H_0^1 \cap H_0^2 \cap H_0^3$			
	$(\alpha_{21} = \alpha_{31} = \alpha_{41} = \alpha_{51} = 0)$					
	1.23		11.18			
	(0.54)		(0.19)			

Note: the superscripts $^{+}$, * and ** indicate rejections at 10%, 5% and 1% significance levels, respectively.

As for the short-run dynamics, we proceed by estimating equations (10) and (11) over the full sample 1980 : Q4–1998 : Q4 using M3 computed using the fixed conversion rates. The results are summarized in equations (12) and (13) below and in Figure 5, which show some graphical evaluation statistics:

$$(\tilde{m} - p)_t = const + \underset{(0.059)}{1.111}\, y_t - \underset{(0.391)}{0.710}\,(RL_t - RS_t) - \underset{(0.370)}{1.201}\,\pi_t + \hat{u}_t \qquad (12)$$

$$\begin{aligned}\Delta(\tilde{m} - p)_t = & -\underset{(11.69)}{0.632} + \underset{(1.82)}{0.074}\,\Delta^2 y_t + \underset{(3.07)}{0.228}\,\Delta_2 RS_t/2 - \underset{(3.95)}{0.320}\,\Delta RL_{t-1} - \underset{(9.71)}{0.478}\,\Delta_2\pi_t/2 \\ & - \underset{(11.81)}{0.126}\,[(\tilde{m} - p) - 1.111y + 0.710(RL - RS) + 1.201\pi]_{t-2} \\ & - \underset{(4.94)}{0.0096}\,DUM86_t + \hat{v}_t \end{aligned} \qquad (13)$$

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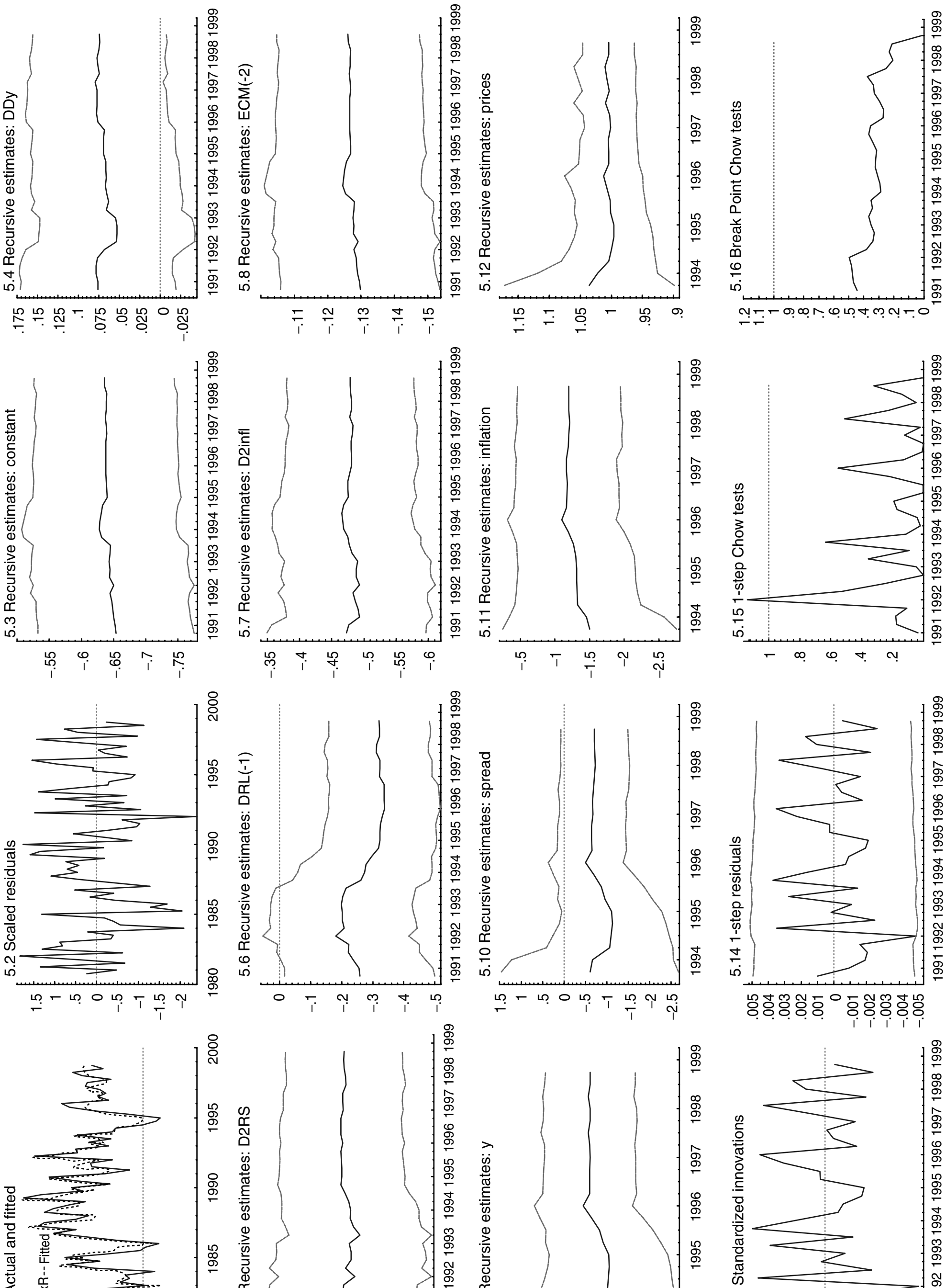

Figure 5. Graphical evaluation and recursive estimates: fixed conversion rate

$T = 73$ (1980 : Q4–1998 : Q4)	$R^2 = 0.80$	$\sigma = 0.235\%$	$DW = 2.03$
$LM(1) = 0.022(0.88)$	$LM(4) = 0.904(0.35)$	$LM(1, 4) = 1.40(0.24)$	
$ARCH(4) = 0.624(0.65)$	$HET = 0.319(0.98)$	$NORM = 0.234(0.89)$	
$RESET = 0.319(0.98)$	$RED = 0.58(0.72)$	$HANS^1 = 0.105$	$HANS^2 = 0.679$

where $\tilde{m}$ denotes that the historical series for M3 is calculated using the fixed conversion rates. As before, equation (12) is obtained as the long-run solution of an ADL(2) model and equation (13) parsimoniously encompasses the unrestricted autoregressive distributed lag model.

Comparison of equations (12) and (13) with (10) and (11) indicates that the change of aggregation method for M3 does not have any noticeable impact either on the long-run or the short-run parameters of money demand, with differences in point estimates always well within one standard error bound.[17] That in turn suggests that it is unlikely that the shift to the calculation of euro area M3 using fixed conversion rates may in itself render our results useless in the context of Stage Three.

Although (based on the respective estimated standard errors) there appears to be little difference between both aggregation procedures, encompassing tests are conducted with a view to discriminating between the two competing models (13) and (11). In order to deal with different dependent variables, we follow the approach that Ericsson, Hendry and Tran (1994) proposed for testing models with seasonal adjusted data versus models with seasonally unadjusted data. In this context, the hypothesis that (11) encompasses (13) (H_0^a) can be tested by adding $\Delta(m - \tilde{m})$ and the error-correction term in (13) to equation (11) and testing for their joint significance. Conversely, the null that (13) encompasses model (11) (H_0^b) can be tested by adding $\Delta(m - \tilde{m})$ and the error-correction term in (11) to equation (13). Likelihood ratio tests for these hypotheses provide the following results: $\chi_2^2 = 1.29$ (0.524) and $\chi_2^2 = 4.25$ (0.120) for the nulls H_0^a and H_0^b respectively. Therefore, although neither H_0^a nor H_0^b can be formally rejected at standard confidence levels, it appears that the evidence for H_0^b is less compelling, somewhat favouring (11) over (13). Recursive computation of the encompassing tests tends to confirm this conclusion.

5. CONCLUSION

Given the prominent role assigned to monetary aggregates in the monetary policy strategy of the ECB, the present paper has analysed some empirical evidence on the demand for broad money M3 in the euro area. Starting with the system results, three long-run relationships with economic content were found: first, the spread between the long-term and the short-term nominal interest rates, consistent with the expectations theory of the term structure of interest rates; second, the long-term and short-term real interest rates, consistent with the Fisher parity; and, third, a long-run demand for broad money M3 in the euro area.

Turning to our central relation of interest, money demand, the empirical findings are plausible from a number of perspectives. From an economic point of view, the long-run demand for M3 in the euro area shows a simple specification whereby holdings of M3 real balances are determined

[17] In particular, concerning the long-run relationship (12), the spread restriction continues to be accepted [$F(1, 57) = 0.041$ (0.840)] while the exclusion of inflation and the unit income elasticity are rejected, though the latter only marginally [$F(1, 58) = 10.53$ (0.002) and $F(1, 58) = 3.48$ (0.067) respectively]. Rejection of the unit income elasticity is consistent with the evidence from the system estimates in Table II [$\chi_3^2 = 9.69$ (0.021), when the cointegration space is estimated unrestrictedly].

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by real income, the opportunity cost of holding money relative to financial assets not included in the definition of the monetary aggregate and the inflation rate. The signs and magnitudes of the estimated long-run coefficients also appear quite plausible on theoretical grounds. The long-run income elasticity is estimated significantly above one (1.13), a finding which is often interpreted as proxying omitted wealth effects in the demand for money function. The estimated long-run semi-elasticities with respect to the spread and the inflation rate are, respectively, -0.87 and -1.51.

In the short run, changes in M3 holdings were modelled via an equilibrium correction model whereby growth in real M3 is explained by the changes in income, interest rates and inflation, on the one hand, and the extent to which M3 differs from its equilibrium level, on the other. Equilibrium correction models have proven successful in modelling the demand for money in a number of EU countries and in the area as a whole during the last decade. From a statistical point of view, the estimated models for money demand show no sign of mis-specification and appear to be a congruent statistical representation of the 'average' process which has driven broad money M3 in the euro area during the sample period under investigation. When compared with the existing empirical evidence on money demand in individual euro area countries and in the euro area as a whole, as surveyed, for instance, in Browne *et al.* (1997), the estimated models fare quite well in terms of stability, fit, residual standard errors, etc. In particular, the estimates presented support the view reached by many previous studies that area-wide money demand equations outperform most national equations.

The demand for broad money M3 equation turns out to be reasonably constant when estimated recursively. This holds in the face of the dramatic economic changes which euro area countries undertook within the estimation period and poses the question of whether the estimated models are of any use for monetary policy purposes in the context of Stage Three. This issue is of utmost relevance, as it has been argued that the adoption of a single monetary policy marks a significant regime shift, which could affect empirical economic relationships estimated on the basis of pre-1999 data. The potential effects of the adoption of a single currency are also argued to be of particular importance at the outset of Stage Three in the financial sector of the economy, thereby affecting the behaviour of monetary aggregates in the euro area. Unfortunately, it is not possible to test explicitly for parameter invariance against such a historical event, since it has no parallel within our estimation sample, but the estimated models constitute useful benchmarks for monitoring structural change in the early stages of Stage Three and for the evaluation of rival models as they become available. Nonetheless, the empirical finding that the area-wide money demand equation has remained constant against a changing background in monetary policy convergence among euro area countries during the sample period under investigation suggests that instabilities in the models should not arise because of the introduction of the single currency itself, as some of the alleged instabilities should likely have shown up in anticipation of EMU.

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