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NON-LINEAR DYNAMICS IN THE EURO AREA DEMAND FOR M1

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In 2006 all ECB publications will feature a motif taken from the €5 banknote.

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Abstract

This paper investigates possible non-linearities in the dynamics of the euro area demand for the narrow aggregate M1. A long-run money demand relationship is firstly estimated over a sample period covering the last three decades. While the parameters of the relationship are jointly stable, there are indications of non-linearity in the residuals of the error-correction model. This non-linearity is explicitly modelled using a fairly general Markov switching error-correction model with satisfactory results. The empirical findings of the paper are consistent with theoretical predictions stemming from "buffer stock" and "target-threshold" models and with analogous empirical evidence for European countries and the US.

JEL classification: E41, C22.

Keywords: Euro area, cointegration, non-linear error correction, demand for money.

Non-technical summary

Linear models embodying error-correction mechanisms have become the standard macroeconometric tool in the empirical literature on money demand. One of the main reasons for their popularity is that they have been able to provide a statistically meaningful representation of the observed sluggishness in the portfolio allocation behaviour of economic agents. Yet, such sluggishness derives from the existence of market rigidities, such as portfolio adjustment costs, which may also translate into non-linearities in the dynamics of adjustment to equilibrium.

The aim of the paper is to investigate and explicitly model non-linearities in the dynamics of the euro area demand for the narrow aggregate M1. Such non-linearities are usually rationalised on the basis of "target-threshold" and "buffer stock" theoretical models (Miller and Orr; 1966, Cuthbertson and Taylor; 1987 and Gandolfi and Lothian; 1983). The rationale behind these models is that, due to adjustment costs, it may not be optimal for agents to re-adjust immediately their asset portfolios after a shock so as to bring their balances back to the target straight away. By contrast, the optimal response may be to let monetary balances fluctuate as a temporary buffer stock until the other assets can be adjusted. Only when the deviations of monetary holdings from the desired levels become relatively large or exceed some specified thresholds, agents engage in those transactions needed to bring their balances back to the target.

Our empirical investigation is based on Krolzig's (1997) two-stage approach to the modelling of cointegrated vector autoregression systems with Markovian regime-shifts. In the first stage, Johansen's multivariate cointegration procedure is applied to a system of variables including real money, output and a short-term interest rate in order to determine the cointegrating rank and estimate the identified long-run money demand relationship. We find that, while the parameters of the relationship are jointly stable over the last three decades, there are indications of non-linearity in the residuals of the error-correction model. In the second stage such non-linearity is directly modelled by means of a regime-dependant error-correction model. Our results show that the dynamic behaviour of M1 - whereby deviations from equilibrium are corrected - varies depending on the prevailing regime of monetary conditions. In particular, the probabilities of being in the

regime in which the error-correction adjustment is faster are typically higher in periods associated with large deviations from equilibrium. By contrast, the probabilities of being in the regime in which the adjustment to equilibrium is slower are usually higher in correspondence with periods of relatively small deviations from equilibrium.

These empirical findings are consistent with theoretical predictions by buffer stock and target-threshold models. In addition, they are consistent with analogous results for several European countries and the US reported in recent empirical contributions (see for instance Escribano, 2004; Lütkepohl et al., 1999; Sarno, 1999; Teräsvirta and Eliasson, 2001; Ordóñez, 2003; Sarno et al., 2003; and Chen and Wu, 2005).

1 Introduction

Linear models embodying error-correction mechanisms have become the standard macroeconometric tool in the empirical literature on money demand (see Sriram, 2001; Duca and van Hoose, 2004). They combine a theoretically-grounded description of the behaviour of money demand in equilibrium with a data-driven specification of the (linear) dynamics of disequilibrium correction in the short-run. One of the main reasons for their popularity is that these models have been able to provide a statistically meaningful representation of the observed sluggishness in the portfolio allocation behaviour of economic agents. Yet, such sluggishness derives from the existence of market rigidities, such as portfolio adjustment costs, which may also translate into non-linearities in the dynamics of adjustment to equilibrium.

Non-linearities in the dynamics of money demand are typically rationalised on the basis of "target-threshold" and "buffer stock" theoretical models.[1] Miller and Orr's (1966) inventory theoretic model of the demand for transaction balances by firms is a representative example of a target-threshold model, while Cuthbertson and Taylor (1987) and Gandolfi and Lothian (1983) fall under the category of buffer stock models. These models start from the observation that, due to shocks of various nature, the monetary holdings of individual agents may depart from their desired or "target" levels. However, in the presence of adjustment costs, it may not be optimal for agents to re-adjust immediately their asset portfolios so as to bring their balances back to the target straight away. By contrast, the optimal response may be to let monetary balances fluctuate as a temporary buffer stock until the other assets can be adjusted. Only when the deviations of monetary holdings from the desired levels become relatively large or exceed some specified thresholds, agents engage in those transactions needed to bring their balances back to the target.

These theoretical frameworks have been developed to explain non-linearities in the money demand behaviour of individual agents. However, the microeconomic frictions arising from portfolio adjustment costs may also result - under certain conditions - in persistent deviations of the aggregate long-run money demand from the equilibrium level and in non-linearities in the short-run monetary dynamics. Bertola and Caballero (1990) argue that, in

[1] For a discussion on the notion of buffer stock in monetary economics see Laidler (1984). Mizen (1994) is a comprehensive study of buffer stock money demand models, also including target-threshold models as a special type.

the presence of kinked adjustment costs, such conditions are related to the degree of coordination and synchronisation across individual agents, which is - in turn - likely to depend on the relative importance of aggregate and idiosyncratic uncertainty. When the former predominates, also the aggregate variables display, at least to some extent, the type of sluggish dynamic adjustment associated with microeconomic money demand frictions.

Consistently with these theoretical predictions, in recent years some authors have found empirical evidence of non-linearities in the short-run dynamics of monetary aggregates. Sarno (1999), Lütkepohl et al. (1999), Teräsvirta and Eliasson (2001), Ordóñez, (2003), Sarno et al. (2003), and Chen and Wu (2005) model such non-linearities for various European countries and the US using regime-dependent models, usually smooth-transition regressions. Hendry and Ericsson (1991) and Escribano (2004) instead model the non-linearities in the money demand in the UK using a cubic polynomial error-correction model.

The purpose of this paper is to investigate whether there is similar evidence of non-linearity in the (short-run) dynamics of the demand for euro area M1. In general terms, the focus on a narrow aggregate such as M1 can be explained by the fact that it is a close empirical counterpart of the notional monetary balances featuring in the relevant theoretical models (e.g. Miller and Orr's (1966) target-threshold model). Besides, like other monetary aggregates, M1 can effectively summarise the information available in key macroeconomic fundamentals, such as output, prices and interest rates. Nelson (2003) has recently noted that, by proxying a spectrum of yields that matter for aggregate demand but are not always directly observable, monetary aggregates such as M1 may provide incremental information about aggregate demand.

There are also factors specific to the euro area that render the analysis of the dynamics of M1 of significant interest for monetary policy purposes. Indeed, in the euro area M1 exhibits a number of empirical properties that make it an important component of the information set available to policy makers.[2] In particular, changes in real M1 seem to contain useful information about developments in area-wide output up to three years ahead. In addition, over the last two decades turning points in M1 growth have often reliably predicted those in the general euro area business cycle with a lead of around three to four quarters. Against this background, an in-depth understanding

[2] See Issing (2003) and the studies and the references therein.

of the dynamics of M1, with a focus on its potential non-linearities, would enhance the information content of this monetary aggregate for future output activity and, ultimately, prices.

In line with much of the quoted empirical literature, this paper characterises non-linearity in terms of regime-dependency in the dynamic behaviour of money, i.e. allowing for the possibility that the short-run dynamics of money demand varies across different states of the economy. However, an innovation of this study is the choice - based on an extensive specification search - of a Markov-switching error-correction model to characterise such regime-dependency.[3] In particular, the study applies Hamilton's (1989) Markov-switching model, as extended to cointegrated vector autoregression models by Krolzig (1997).[4] The model is estimated over a sample period covering the last three decades. To our knowledge, this is the first money demand study for the euro area estimated over such extended sample.

Consistently with theoretical predictions by buffer stock and target-threshold models and with previous empirical results for the US and some European countries, we find that the error-correction model of real euro area M1 is characterised by non-linear dynamics of adjustment to monetary disequilibria. In particular, when the deviations of aggregate demand for monetary balances from equilibrium are large, the speed of adjustment to the desired level of monetary balances is faster.

2 The long-run money demand relationship

Our empirical investigation relies on Krolzig's (1997) two-stage approach to the cointegration analysis of vector autoregression (VAR) models with Markovian regime-shifts.[5] In the first stage (which is the object of this section), Johansen's (1995) multivariate cointegration procedure is applied to a system of variables in order to determine the cointegrating rank and estimate

[3] A number of specifications for smooth-transition models (mainly single equations), were also tested, but it was not possible to estimate with precision the parameters governing the regime transition.

[4] Camacho (2005) has recently proposed an alternative model of Markov-switching equilibrium adjustment based on a common trends representation.

[5] The empirical results have been obtained using the packages Ox, PcGive and MSVAR (downloadable from H.-M. Krolzig's web page at http://www.kent.ac.uk/economics/staff/hmk/index.htm)

the identified long-run money demand relationship.[6] In the second stage, a Markov-switching model of the dynamics of monetary balances is selected and estimated, conditional on the previously obtained cointegrating matrix.

The analysis is based on quarterly data for the euro area – defined according to the principle of changing composition (the 11 original countries up to 2000Q4; these plus Greece, thereafter) - over the period 1971Q4 to 2003Q3. The variables modelled consist of the monetary aggregate M1 (M_t) deflated by the GDP deflator (P_t), real GDP (Y_t) and the short-term market interest rate (R_t). Nominal M1 is the period average of the end-of-month seasonally-adjusted (s.a.) notional stock compiled by the ECB. The GDP data are based on the aggregation of s.a. national accounts data (ESA95 whenever available) up to 1998Q4; hereafter, on area-wide Eurostat statistics. The national data on M1 and GDP prior to the introduction of the euro have been aggregated using the irrevocable conversion rates announced on 31 December 1998 (19 June 2000 for Greece). The interest rate is a weighted average (based on GDP weights at 2002 purchasing power parities) of national 3-month interbank interest rates up to 1998Q4; thereafter, it corresponds to the three-month EURIBOR.

The long-run money demand function is specified in the following log-log form:

$$(m-p)_t = \beta_1 y_t - \beta_2 r_t + k \tag{1}$$

where all variables are in natural logarithms and k denotes an intercept unrestricted to the cointegrating space. As noted by Lucas (2000) for the US, this functional form presents significant advantages over alternative specifications in terms of sounder micro-foundations and a more accurate calculation of the welfare costs of inflation at low interest rates. In addition, in the framework of the shopping-time model of money demand determination by McCallum and Goodfriend (1987), Lucas observes that, for reasonable estimates of the interest rate elasticity, the log-log specification is more in line with theoretical models, such as Miller and Orr (1966).[7] For the euro area,

[6] Note that in the first stage it is not necessary to model the Markovian regime shifts explicitly in order to derive the equilibrium relationships (Saikkonen, 1992).

[7] Chadha et al. (1998) concur on the theoretical superiority of the log-log form. Based on McCallum and Goodfriend's (1987) model, they show that the choice of any well-behaved utility function and transactions technology (e.g. Cobb-Douglas, CES and translog functions) is likely to result in a log-log specification of long-run money demand. However, using UK data, they find that the empirical advantages of the log-log specification may be more relevant for the short-run dynamics of the money demand than for its

Stracca (2003) investigates the issue of the choice of the functional form for the long-run demand for M1, providing empirical evidence in support of the log-log specification.

As a preliminary step, the statistical properties of the variables forming the system $z = [(m-p), y, r]$ are examined using standard unit root tests (augmented Dickey-Fuller and Phillips-Perron) as well as the KPSS stationarity test. The results - not reported for the sake of brevity - suggest that over the sample period considered all the variables should be modelled as I(1) in levels.

The cointegrating properties of the system z_t are subsequently tested by means of the multivariate cointegration procedure by Johansen (1995):

$$\Delta z_t = v + \sum_{i=1}^{p-1} \Gamma_i \Delta z_{t-i} + \Pi z_{t-1} + \Psi D_t + u_t \tag{2}$$

where the parameters of the model are represented by the vector v of deterministic components, the matrices Γ and Ψ of short-run coefficients, and the matrix $\Pi = \alpha\beta'$, with α the vector of loading factors and β the matrix of long-run coefficients. In particular, $\beta' z_{t-1}$ includes the one-period lagged money-demand error correction term implied by the cointegrating vector, D_t is a vector of I(0) exogenous variables and u_t is the errors vector (assumed to be serially non-correlated with zero mean and constant covariance matrix). Consistently with Stracca (2003), D_t includes two impulse dummies ($ID99Q1$ and $ID00Q1$) taking the value 1 in the first quarter of 1999 and 2000, respectively, and zero elsewhere, as exogenous variables.[8]

The application of the Johansen (1995) procedure enables us to determine the number of cointegrating vectors and, subject to appropriate specification testing, allows to identify and estimate such vectors. On the basis of the Akaike, Hanna-Quinn and Schwartz information criteria, the lag order p of

equilibrium behaviour.

[8] The first dummy is introduced in order to control for the exceptionally large rise in the demand for M1 holdings (especially for overnight deposits) recorded after the start of Stage Three of European Monetary Union in January 1999. This rise probably reflected institutional innovations associated with the new monetary policy regime (e.g. the introduction of a new reserve requirements system) as well as the changes in statistical reporting procedures. The second dummy controls for the temporary acceleration in demand for currency at the time of the "millennium bug" scare, when concerns about possible disruptions to retail payment systems and cash dispensing machines became widespread in several euro area countries.

the testing VAR (including linear trends in the data and an unrestricted intercept in the cointegrating vector) is set at 2 in levels. Panel A of Table 1 reports the Johansen's trace (λ_{trace}) and maximum eigenvalue ($\lambda_{\max}$) cointegrating tests. Both tests reject the hypothesis of no cointegration at the conventional significance levels, while accepting that of at most one cointegrating relationship. The evidence of cointegration is robust to the use of test statistics adjusted for degrees of freedom (as suggested in Reimers, 1992) in order to control for potential small-sample bias.

The results of the long-run exclusion tests in Panel B show that none of the variables can be excluded from the cointegrating vector at the conventional significance levels. Furthermore, the tests for weak exogeneity reveal that y and r can be treated as weakly exogenous to the system, both individually and jointly.

The estimated cointegrating vector, normalised with respect to real M1 and to zero mean, is also presented in Table 1. From Panel C it is possible to see that the estimated income elasticity is 0.744. This value is consistent with theoretical predictions as it falls between the value of 0.5 anticipated by the Baumol-Tobin inventory-theoretic model of transaction demand for money and the unitary elasticity implied by the quantity theory.[9] The interest rate elasticity of the demand for real M1 is estimated at –0.392. Because of the relatively low and sluggish average remuneration of the deposits included in M1 (which also includes zero-remunerated currency in circulation), this interest rate can be interpreted as approximating the opportunity cost of holding the monetary aggregate. Given the functional log-log form, the interest rate elasticity is constant across interest rates and measures the percentage change in the demand for money in response to a one percent change in the short-term interest rate.[10] On the basis of the magnitude and sign of the coefficients, this cointegrating vector can be interpreted as representing a long-run demand function for real M1.

Given the relatively broad time span covered by the sample period, which comprises periods of both high and low interest rates, it is important to test for the stability of the coefficients of the equilibrium money demand relation-

[9] In the conditional model, the hypothesis of a unitary income elasticity is rejected by the data (χ^2_1=8.70 [p-value=0.04]).

[10] The value of the interest rate elasticity is consistent with the finding of Stracca (2003) for euro area M1, which reports a coefficients of 0.51, and with the calibration of the US money demand model by Lucas (2000), which sets the value of the elasticity at 0.50.

ship. For this purpose, we apply two types of Nyblom tests for parameter constancy of the cointegrating vector as extended to cointegrated VARs by Hansen and Johansen (1999). The null hypothesis of the tests - which are respectively based on the maximum (Sup) and the mean (Mean) of a weighted LM-type statistics over the sample period - is the joint stability of the parameters of the cointegrating vector. The supremum and mean test statistics yield 1.60 [p-value=0.53] and 0.98 [p-value=0.20], respectively.[11] The high level of the p-values indicates that the null hypothesis cannot be rejected at the conventional significance levels, suggesting that the long-run parameters are jointly-stable over a sample period covering the last three decades.[12]

Conditional on the finding of joint weak exogeneity for y and r, the dynamic model is specified as a single equation error-correction model. The estimated equation is reported in Panel D. In particular, the coefficient of the error correction term is negative and statistically significant, supporting the interpretation of the cointegrating vector as a long-run money demand function. Yet, the relatively small size of the coefficient (-0.051) reveals a rather sluggish adjustment to equilibrium in case of deviations. This is illustrated in Figure 1 by the slow rate at which monetary disequilibria are corrected.

Finally, the statistical properties of the residuals of the model are evaluated by means of several standard misspecification tests for autocorrelation, non-normality and heteroscedasticity. The results are satisfactory and suggest that the model is adequately specified. However, we fail to reject the null-hypothesis of no mis-specification of the RESET test. Originally developed to test for omitted regressors, a significant value of the RESET statistic is often indicative of non-linearity in the residuals (see Granger and Teräsvirta, 1993). The evidence of mis-specification provided by this test suggests that the specification of the equation may be improved by modelling explicitly such non-linearity. The next section formally investigates this issue.

[11] The distributions of the tests are bootstrapped using 1,000 replications. The computations are performed using the program Structural VAR, version 0.19, by Anders Warne (downloadable from www.texlips.hypermart.net/svar).

[12] However, it should be noted, as a caveat, that the Nyblom test assumes a linear short-run dynamics.

3 Modelling the non-linear dynamics of M1

The analysis of the residuals of the linear error-correction model suggests that a standard model with time invariant parameters may not provide an appropriate representation of the short-run dynamics of M1. Such dynamics may be better captured by a model allowing for some form of regime-dependent behaviour. In particular, if the non-linear process is time-invariant conditional on an (unobservable) regime variable s_t, a Markov-switching model may be considered as an appropriate framework. The idea behind this class of model is that the parameters of the underlying data generating process of the observed time series vector z_t depend upon an unobservable regime variable s_t, representing the probability of being in a certain state of the world.

Letting $s_t \in \{1, ...M\}$ indicate the regime prevailing at time t, the properties of the MS(M)-ECM(p) model for the euro area real money demand can be analysed depending on the realization of the regime:

$$\Delta(m-p)_t = v\left(s_t\right)+\gamma_0\left(s_t\right)\Delta x_t+\sum_{i=1}^{p-1}\gamma_i\left(s_t\right)\Delta z_{t-i}+\alpha\left(s_t\right)\beta' z_{t-1}+\psi\left(s_t\right)D_t+u_t \tag{3}$$

where $v\left(s_t\right)$ is the regime-dependent intercept term, $\gamma_0\left(s_t\right)$, $\gamma_i\left(s_t\right)$ and $\psi\left(s_t\right)$ are the vectors of short-run parameters, $\alpha\left(s_t\right)$ is the state-dependent adjustment coefficient, $x = [y, r]$, $\beta' z_{t-1}$ is the cointegrating vector and D_t is the vector including the two dummies. The hypothesis underlying the model is that the equilibrium relationship does not vary across regimes; it is only the speed of the adjustment to the error-correction term that is allowed to vary. Finally, note that the error term u_t also depends on the realisation of the regime since $u_t \sim NID\left(0, \Sigma\left(s_t\right)\right)$.

Since parameters depend on a regime which is assumed to be stochastic and unobservable, a generating process for the states s_t has to be formulated. In particular, the stochastic process generating the unobservable regimes is assumed to be an ergodic Markov chain defined by the transition probabilities:

$$p_{ij} = \Pr\left(s_{t+1} = j \mid s_t = i\right), \qquad \sum_{j=1}^{M} p_{ij} = 1 \qquad \forall i, j \in \{1, ..., M\}. \tag{4}$$

By inferring the probabilities of the unobservable regimes conditional on the available information set, it is possible to reconstruct the evolution of the regimes.

In order to select the best specification of the model for the euro area data we run a battery of tests. We start with linearity tests against the various types of Markov-switching models with two regimes and subsequently use different statistics to select the best among the possible Markov-switching specifications. The first column of Table 2 reports the p-value of the (upper-bound of the) Likelihood-Ratio (LR) statistic testing the null-hypothesis of linearity against the alternative of a specific type of Markov-switching non-linearity.[13]

On the basis of the LR test, the data fail to reject the null of linearity for the models specifying the regime switching behaviour for either the intercept term (MSI), or the short-run parameters of the error-correction model (MSA), or the variance covariance matrix (MSH). By contrast, the null of linearity is easily rejected at the conventional significance levels for the specifications combining different types of regime-dependence behaviour: in the intercept and short-run coefficients (MSIA), in the intercept and variance-covariance matrix (MSIH) and in the intercept, short-term parameters and variance-covariance matrix (MSIAH). Only for the MSAH model, which combines regime dependency in the short-run parameters and the variance-covariance matrix, the null of linearity cannot be rejected. These results suggest that in order to identify and describe the regimes it is necessary to use models specifying general forms of regime-switching dynamics, such as the MSIA, MSIH or MSIAH models.

The second column of Table 2 shows the p-values of LR restriction tests designed to select the most parsimonious among the candidate Markov-switching specifications. In practice, these tests assess each specification against the more general MSIAH model. Based on the results of the non-linearity LR tests, we restrict the discussion to the last three specifications. The null hypothesis of no shifting in the short-run parameters (MSIH versus MSIAH) is strongly rejected by the data. By contrast, the null of no shifting in the variance-covariance matrix (MSIA versus MSIAH) cannot be rejected.

On the basis of this test, the MSIA specification presents some advantages over the less parsimonious MSIAH model. However, there are some indica-

[13] The application of LR tests in the context of Markov-switching models is discussed in Hansen (1992, 1996).

tions that the MSIAH specification is to be preferred. First, the Regime Classification Measure (RCM) proposed by Ang and Bekaert (2002) to discriminate among different types of Markov-switching models (third column of Table 2) suggests a better fit of the MSIAH. The RCM is a summary point statistic of the degree of accuracy with which a model identifies the regime switching behaviour over the sample period. The statistic ranges between 0 and 100, with 0 denoting a perfect regime classification performance and 100 indicating that the model fails to provide any information on the regime-dependence. The value of the RCM statistic recorded for the MSIAH specification is fairly low, and significantly smaller than that for the MSIA. In addition, the MSIAH model seems to have an higher explanatory power as can be evinced from the larger value of the coefficient of determination (adjusted for degrees of freedom): 0.66 versus 0.63 (fourth column of Table 2). Finally, the dating cycle identified by the MSIA model is relatively volatile and hard to relate to economic developments in the euro area over the sample period.

On the basis of the above considerations, we restrict our attention to the MSIAH specification. As a final check, we test whether it may be statistically more appropriate to use a model allowing for 3 instead of 2 regimes (fifth column of Table 2). The results of the test are not clear-cut. The null of a two-regime MSIAH model versus a three-regime model can be rejected only at the 10% significance level. However, given the size of the sample (128 observations), we retain the specification allowing for fewer regimes.

The results for the estimation of the MSIAH(2)-ECM(1) model for the euro area (real) M1 are presented in Table 3. The number of observations in each regime is large enough to allow for robust statistical inference.[14] The regimes are fairly persistent, with the conditional probabilities ($p_{11} = 0.94$, $p_{22} = 0.90$) implying an expected duration of around $4\frac{1}{2}$ years and $2\frac{1}{2}$ years for the first and second regime, respectively.

Standard misspecification tests (not reported for the sake of brevity) fail to reveal signs of autocorrelation, non-normality or heteroscedasticity for both the standardised residuals and the one-step prediction errors, suggesting that the model is satisfactorily specified.[15]

[14]Note that the two impulse dummies both fall under regime 1. Consistently with this, their impact on real monetary growth is statistically significant only in first regime (see Table 3).

[15]However, the results of these tests should be interpreted with caution given that their asymptotic distributions may not be valid for residuals from Markov-switching models.

Figure 2 depicts the smoothed probabilities of being in Regime 1 together with the annual growth rate of real M1. Regime 1 includes the periods of highest volatility in real monetary growth over the last thirty years. In particular, it comprises a protracted period of relatively low but volatile growth in real M1 throughout most of the 1970s and beginning of the 1980s as well as a long time span of relatively high and volatile monetary growth throughout the 1990s. By contrast, the probabilities of Regime 2 are associated with periods of more stable money growth.

More formal evidence in support of this observation is provided by the analysis of real money's regime-dependent variances using Warne's (1998) probability weighted estimator of conditional moments.[16] The conditional variance for real M1 growth in Regime 1 (0.016%) is indeed more than twice that in Regime 2 (0.006%).

The theoretical models surveyed in the introductory section lead to the prediction that the process of adjustment to equilibrium should be more effective during the first regime - characterised by more extreme developments in monetary balances - than in the second one. Indeed, buffer stock models would suggest that in periods when the behaviour of money deviates significantly from its norm, agents should adjust to the "desired" level at a higher speed than in tranquil periods. The regime-dependent coefficient of adjustment provide some support to this hypothesis. In both regimes the coefficients of adjustment have the expected negative sign and are significantly different from zero. However, in Regime 1 the estimated coefficient is larger in absolute terms than in Regime 2 (0.073 and 0.053, respectively), confirming the hypothesis that the differences in the speed of monetary disequilibria adjustment depend on the prevailing monetary conditions.[17] While the value of the coefficient of adjustment in Regime 2 is fairly close to the estimate for the linear model, the estimated loading factor in Regime 1 implies a faster correction to the equilibrium. Ceteris paribus, the process of monetary disequilibrium adjustment should be about $1\frac{1}{4}$ years shorter in the first regime than in the second one.

[16] Warne (1998) proposes to compute the conditional moments of a variable by weighting the observed data with the estimated smoothed regime probabilities.

[17] Based on a Wald test, the null hypothesis that the coefficient of adjustment of the conditional model for Regime 1 equals that of the Regime 2 model could be rejected at the 10% signficance level (χ^2_1=3.13 [p-value=0.08]).

These stylized facts find further confirmation in the behaviour of the error-correction term (see Figure 1). The probability-weighted conditional variance of the error-correction term is noticeably higher in Regime 1 (1.13%) than in Regime 2 (0.50%), reflecting the concentration of large disequilibria in the former. High probabilities of being in Regime 1 - the regime in which the coefficient of adjustment of the error-correction term is higher - are typically associated with periods in which the deviations from equilibrium are large. By contrast, the probabilities of being in Regime 2 - in which the adjustment to equilibrium is slower - are usually higher in correspondence with periods of relatively small deviations from equilibrium.

To sum up, our empirical findings provide evidence of non-linearities in the dynamic behaviour of the demand for euro area M1. These findings are consistent with the theoretical predictions of buffer stock and target-threshold models that postulate frictional adjustment in individual money demand behaviour. Our findings are also consistent with the results of studies for various countries (both within and outside the euro area), based on alternative specifications of non-linear error-correction models.

4 Concluding remarks

The empirical analysis presented in this paper supports the use of M1 as an information variable for the conduct of monetary policy in the euro area. Using a log-log functional form, we find evidence of the existence of a stable equilibrium relationship linking the demand for M1 with output, prices and interest rates over a sample period comprising the last thirty years. To our knowledge, this is the first euro area money demand study using such extended sample period. Given the switch to a regime of low and stable inflation within the sample period, it is interesting to note that a formal Nyblom test of parameter constancy indicates that the estimated equilibrium relationship is fairly stable. More generally, the empirical investigation in this study provides further support to Lucas' (2000) arguments in favour of the use of log-log functional forms to specify the long-run behaviour of money demand.

The stability of the estimated relationship suggests that it may provide an adequate benchmark against which to assess actual movements in M1. Large and persistent deviations of monetary balances from the equilibrium level implied by the estimated relationship may reveal the emergence of potential pressures on future economic activity and, ultimately, prices. More generally, periods of excessively fast or slow monetary growth compared to that predicted by the model may signal the build-up of imbalances in asset markets and balance sheets that may lead to macroeconomic instability (Borio and Lowe, 2002).

Based on a fairly general Markov-switching error-correction model, our empirical investigation reveals evidence of non-linearities in the behaviour of the short-run demand for euro area real M1. We find that the dynamic behaviour of M1 - whereby deviations from equilibrium are corrected - varies depending on the prevailing regime of monetary conditions. In particular, the probabilities of being in the regime in which the error-correction adjustment is faster are typically higher in periods associated with large deviations from equilibrium. By contrast, the probabilities of being in the regime in which the adjustment to equilibrium is slower are usually higher in correspondence with periods of relatively small deviations from equilibrium.

Our empirical findings of non-linearity in the dynamics of euro area money demand are consistent with theoretical predictions by buffer stock and target-threshold models. In addition, they are consistent with analogous results for several European countries and the US reported in recent empirical contributions (see Sarno, 1999; Teräsvirta and Eliasson, 2001; Ordóñez, 2003; Sarno et al., 2003; and Chen and Wu, 2005). The fact that these studies have used alternative types of non-linear error-correction models (typically, cubic polynomial or smooth-transition models), suggests that the finding that frictions in individual money demand behaviour translate into rigidities at the aggregate level may be fairly robust to the choice of econometric methodology.

One potential implication of our findings of non-linearities in euro area monetary dynamics is that the effects of excessively fast or slow monetary growth on the economy could also be regime-dependent. This would imply that the assessment of the implications for output of monetary imbalances should be preceded by an accurate analysis of the monetary conditions characterising the state of the economy. A failure to do so may lead to an inappropriate interpretation of the information contained in monetary developments.

Future work should aim to establish whether similar asymmetries can be identified also in the dynamics of weighted monetary aggregates, such as the Divisia aggregates recently extended to a monetary union set-up by Barnett (2003).

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Table 1 Johansen procedure

A. Cointegration tests

Eigenvalues	Rank	λ_{trace}	$\lambda^{\dagger}_{trace}$	$\lambda_{\max}$	$\lambda^{\dagger}_{\max}$
0.22109	$=0$	45.46**	43.33**	31.98**	30.48**
0.09891	≤ 1	13.47	12.84	13.33	12.71
0.00112	≤ 2	0.14	0.14	0.14	0.14

B. χ^2 restriction tests (conditional on unitary rank)

	$(m-p)$	y	r
Exclusion	χ^2_1 =17.54 [0.00]	χ^2_1 =13.00 [0.00]	χ^2_1 =18.00 [0.00]
Weak exogeneity	χ^2_1 =13.44 [0.00]	χ^2_1 =2.43 [0.12]	χ^2_1 =1.80 [0.18]
Joint weak exogeneity (y and r)		χ^2_2 =2.99 [0.22]	

C. Estimated cointegrating vector (conditional on weak exogeneity of y and r)

$$(m-p) = \underset{(0.07)}{0.744}y - \underset{(0.04)}{0.392}r$$

D. Dynamic money demand equation

$$\Delta(m-p)_t = \underset{(0.001)}{0.005} - \underset{(0.009)}{0.051}ECT_{t-1} + \underset{(0.073)}{0.178}\Delta(m-p)_{t-1} - \underset{(0.118)}{0.084}\Delta y_t$$
$$+\underset{(0.122)}{0.013}\Delta y_{t-1} - \underset{(0.009)}{0.024}\Delta r_t - \underset{(0.009)}{0.022}\Delta r_{t-1} + \underset{(0.007)}{0.027}ID99Q1_t$$
$$+\underset{(0.007)}{0.031}ID00Q1_t + \varepsilon_t$$

$$T = 128; AdjR^2 = 0.65; s.e.(\varepsilon_t) = 0.68\%; LM(1) : F(1,118) = 1.52[0.22];$$
$$LM(1-5) : F(5,114) = 0.79[0.56]; ARCH(1-4) : F(4,111) = 0.63[0.64];$$
$$NORM : \chi^2_2 = 1.58[0.45]; HET : F(14,104) = 1.21[0.28];$$
$$RESET : F(1,118) = 6.01[0.02]$$

Note: † denotes adjustment for degrees of freedom as in Reimers (1992);
** (*) rejection of the null hypothesis at 1% (5%) critical level;
P-values in square brackets; standard errors in parentheses.

Table 2 Identification procedure

	LR-linearity	LR-restrictions	RCM	Adj R^2	LR-regimes
MSI	0.143	0.000	36.6	0.64	0.182
MSA	0.367	0.179	63.2	0.65	0.013
MSH	0.999	0.000	99.8	0.62	0.999
MSAH	0.294	0.408	18.5	0.65	0.951
MSIA	0.048	0.981	31.3	0.63	0.132
MSIH	0.047	0.002	25.7	0.64	0.728
MSIAH	0.034	-	21.7	0.66	0.093

Note: For the LR tests (see Hansen, 1992, 1996) only p-values are reported.
LR-linearity is a test of the null hypothesis of linearity against each possible Markov-switching specification.
LR-restrictions tests each Markov-switching specification against the more general MSIAH model.
RCM is the Regime Classification Measure (RCM) by Ang and Bekaert (2002).
LR-regimes is a test of the null hypothesis of 2- versus 3-regimes.

Table 3 MSIAH(2)-ECT(1) estimation for $\Delta(m-p)_t$

Transition probabilities			**Regime properties**		
	Reg 1	Reg 2		nObs	Duration
Reg 1	0.9429	0.0571	Reg 1	80.5	17.5
Reg 2	0.0971	0.9029	Reg 2	47.5	10.3

Regime 1			**Regime 2**		
	Coef	s.e.		Coef	s.e.
$Const$	0.004	0.001	$Const$	0.009	0.002
$\Delta(m-p)_{t-1}$	-0.036	0.094	$\Delta(m-p)_{t-1}$	0.267	0.130
Δy_t	-0.310	0.165	Δy_t	0.024	0.129
Δy_{t-1}	0.089	0.159	Δy_{t-1}	-0.243	0.124
Δr_t	-0.028	0.011	Δr_t	-0.005	0.009
Δr_{t-1}	-0.021	0.013	Δr_{t-1}	-0.024	0.012
ECT_{t-1}	-0.073	0.012	ECT_{t-1}	-0.053	0.014
$ID00Q1$	0.033	0.007	$ID00Q1$	0.027	0.039
$ID99Q1$	0.031	0.007	$ID99Q1$	0.022	0.029
Std error (Reg.1)		0.62%	Std error (Reg.2)		0.41%

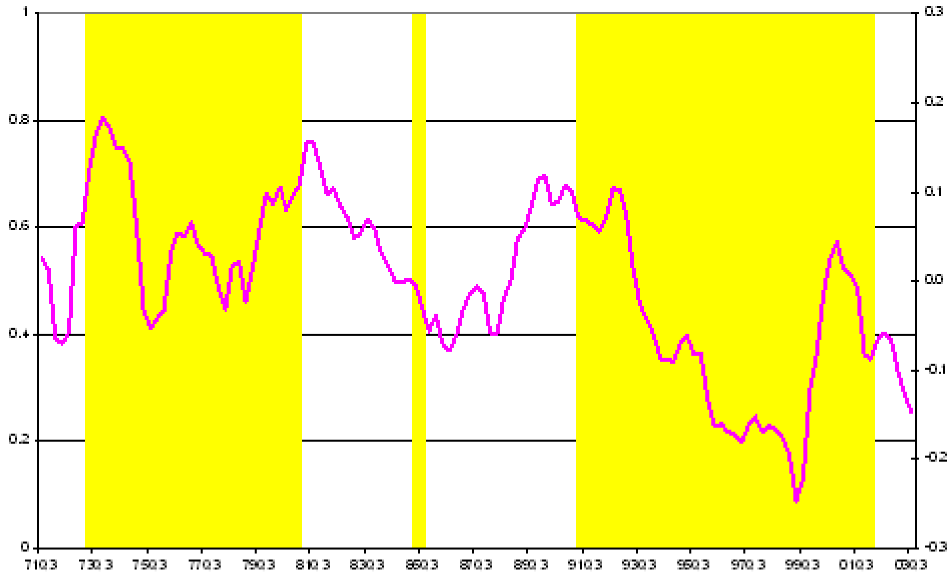

Figure 1: Money demand error-correction term and the regime switching

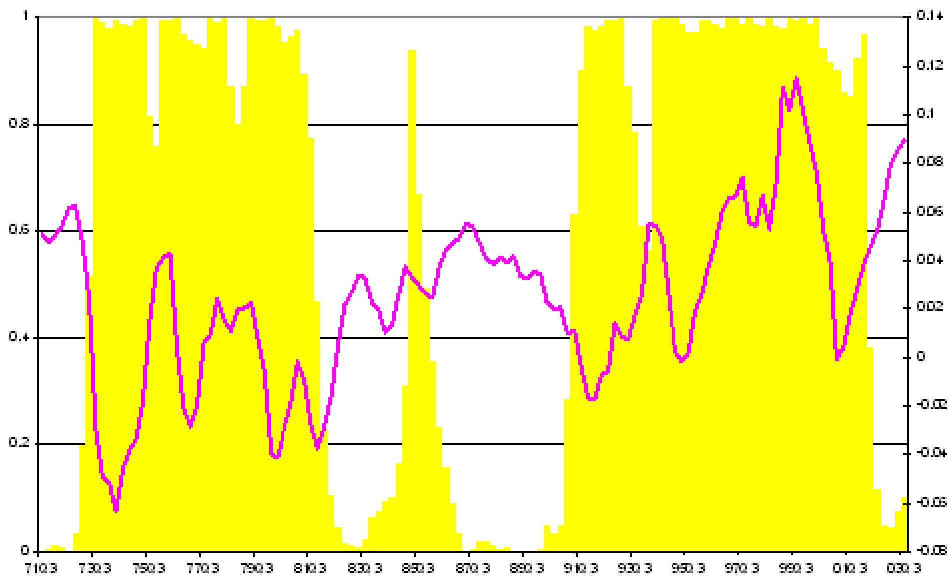

Figure 2: Real money growth and Regime 1 smoothed probabilities

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