

Demand for Broad Money in Italy: Empirical Results for 1972 - 1998*

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Abstract

A money demand function for M2 is estimated for Italy for the period 1972-1998 within an error correction framework. This period has been characterized by major structural changes in the Italian financial system and by major changes in monetary policy. This study takes these changes into account. Moreover, currency substitution, especially between Italy and Germany is incorporated into the model. By accounting for structural breaks and currency substitution a stable money demand function can be found.

Key words: Money demand; error correction model; structural changes

JEL classification: E41, C22

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1 Introduction

The primary goal of the European Central Bank is to maintain price stability. One of the 'two pillars' for achieving this goal is the control of the development of broad money M3 in the euro area. Therefore many studies exist concerning the empirical implementation of money demand for M3 in the euro area, see e.g. COENEN AND VEGA (1999) or HAYO (1999).

But all these empirical work has to use pre-EMU data from the different countries now forming 'Euroland'. The implicit assumption then is that past data give reliable information for monetary policy in EMU, that is, there exists a behavioral persistence. To see, whether this is really the case, it is important to analyze also money demand functions for the individual countries for the pre-EMU period. Therefore we will investigate money demand for Italy, one of the larger countries, and what is especially interesting in this case, with a lot of institutional changes in the pre-EMU era.

During the time from 1972 to 1998 Italian monetary policy and the Italian financial system underwent major structural changes.¹

The seventies were characterized by high inflation and large budget deficits. With the Total Domestic Credit (CTI) as intermediate target economic growth and employment were promoted rather than the stability of prices and of the currency. Italy was isolated from foreign money markets due to extensive controls on capital movements and foreign exchange transactions. In addition, bank lending was regulated by portfolio constraints and credit ceilings.

At the beginning of the eighties the Banca d'Italia started to gain monetary policy independence from the fiscal authorities. A decisive event in this respect was the so called 'divorce' of the Banca d'Italia and the Treasury in July 1981, which freed the central bank from the obligation to intervene and act as residual buyer at the government securities auctions. From 1992 on the choice of the discount rate became the sole decision of the central bank

¹See e.g. FRATIANNI AND SPINELLI (1997).

and the Treasury was no longer allowed to borrow from the Banca d'Italia.

A further important point is the switch from CTI to M2 as intermediate target, which took place between 1983 and 1984 and documented the aim to implement a long-run monetary policy with the main focus on price stability. During the second half of the eighties the quantitative constraints were gradually lowered and the monetary isolation from abroad was removed. Also during that time a lot of innovations concerning financial products and markets were introduced, which caused structural changes in the money market.

Besides the internal challenges Italian monetary policy had to face external constraints arising from its partnership in the European Monetary System (EMS). Italy was in the EMS from the beginning in 1979, but with a wider band than the other members ($\pm 6\%$). In 1990 the band was reduced to the 'standard' band of $\pm 2.25\%$. Before 1983 the Exchange Rate Mechanism was rather flexible with frequent realignments. After 1983 the regime became quasi fixed. Due to credibility problems, a loss of competitiveness and the inability to lower the budget deficit the Italian Lira left the ERM in September 1992, rejoining it in November 1996, this time with the new very wide band of $\pm 15\%$ introduced in August 1993. Italy is now also a member of 'Euroland'.

Although the transition took place over an extended period the year 1983 works well as break point between the two fairly different regimes. Some studies concerning the Italian money demand split their samples at this point and examine two different periods (see for example ANGELINI ET AL. (1994) or JUSELIUS (1998)) or take the year 1983 as beginning of the sample (FANELLI AND PARUOLO (1999) or RINALDI AND TEDESCHI (1996)).

The existing literature concerning the Italian money demand contains many different approaches. SARNO (1999) and MUSCATELLI AND SPINELLI (1996) use historical annual data covering the period from 1861 to 1991, 1990 respectively. Working with single-equation estimation methods they detect one cointegrating relationship. Also applying a single equation model, but for the samples 1975-1979 and 1983-1991 and using monthly as well as quarterly

data, is the work of ANGELINI ET AL. (1994). They estimate a money demand function by allowing for different scale variables (net financial assets and domestic demand) within the two subperiods.

Another approach is used by GENNARI (1999), BAGLIANO (1996), RINALDI AND TEDESCHI (1996) and JUSELIUS (1998), who assume that more than one cointegration relationship exists and therefore implement a multivariate framework. They identify three (BAGLIANO (1996) finds two) cointegrating vectors, one of them being the money demand relationship.

What has not been taken into consideration by the previous literature is the effect of currency substitution. That effects of currency substitution should not be neglected has been made evident for instance by SEITZ AND REIMERS (1999) or DEUTSCHE BUNDESBANK (1995). They show that there are significant substitution effects between the Italian Lira and the DM.

In this paper we are concerned with a structural demand function for broad money. In contrast to the existing literature we

- (i) use data for the whole period from 1972 to 1998, trying to find explanatory variables for the structural changes within this time span, so that the sample must not be split and,
- (ii) since currency substitution has not been taken into account so far, we will test for its presence through US Dollar/Lira and DM/Lira exchange rates and a German interest rate.

This paper is organized as follows. Section 2 contains a preliminary analysis of the data. In section 3 we specify and estimate a long-run money demand relation and use this in an error correction model for real money balances. Stability tests show that it is possible to find a stable money demand equation for the whole period from 1972(1) to 1998(3). Section 4 presents conclusions.

2 The Data and Preliminary Analysis

In the money demand literature there is a general agreement that a long run money demand relation for real balances should include a scale variable such as real gross national product (GNP); this is motivated by the fact that money is used for transaction purposes. Moreover, the rate of interest and inflation are used as explanatory variables to capture opportunity costs of holding money. For a broad monetary aggregate it is also important to include the own rate of return on money. As to Italy, also currency substitution seems to have been significant, see e.g. SEITZ AND REIMERS (1999). Currency substitution has been modelled² through foreign interest rates, the forward discount, the rate of change of the nominal exchange rate and the nominal exchange rate.

Nominal money balances $M2^3$, the Italian and German bond yields are taken from IMF⁴, the own rate of M2 and GNP at current prices are from Banca d'Italia. The GDP deflator P , the Lira/US Dollar exchange rate and Lira/DM exchange rate have been taken from the OECD⁵. All data with the exception of P are seasonally unadjusted. All variables are in logarithms with the exception of the interest rates. m_t denotes the logarithm of the real money stock $M2$ at time t , y_t is the logarithm of real GNP, r_t (r_t^D) is the long-term Italian (German) interest rate and r_t^o the own rate on $M2$ in decimals. The annual inflation rate π_t is defined as the fourths differences of $\log P$ also in decimals. The logarithm of the Lira/US Dollar (Lira/DM) exchange rate is denoted as e_t^{US} (e_t^D). The estimation period ranges from 1972(1) to 1998(3). The variables are depicted in *Figure 1*. m_t shows the most unstable behaviour. There seems to be a break in the data around 1983, which could be associated to the structural changes described above. Moreover a further break is observed in 1995/1996, which could be seen as a

²See TULLIO ET AL. (1996).

³ $M2$ contains notes and coins, current account deposits, postal savings deposits, savings deposits and bank certificates of deposits for terms of less than 18 months.

⁴IMF: 136 39M, 136 61, and 134 61, respectively.

⁵OECD, national accounts: 161051 KSA, 167005D, and 127005D, respectively. The exchange rates are expected values, that means that the series are shifted forward for one quarter.

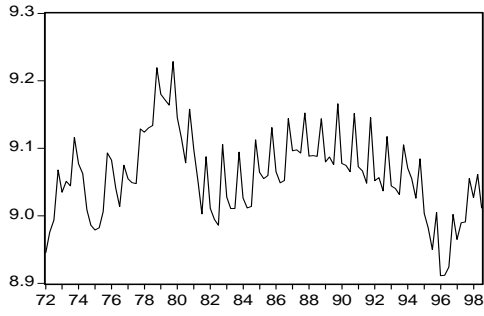
result of the depreciation of the Lira at the beginning of 1995 (DIW (2000)). For y_t the seasonal pattern has changed after 1994. The inflation rate π_t very clearly shows the Italian success in fighting inflation to meet the Maastricht criteria. The exchange rate e_t^D shows obviously that Italy left the ERM in 1992(3).

Applying augmented Dickey-Fuller unit root tests as well as Perron unit root tests allowing for breaks in the constant and/or the trend parameter leads to the result that all time series can be treated as integrated of order one, $I(1)$.⁶ That means the levels of the variables show nonstationary behaviour, but the first differences are stationary. Therefore, the next step in our analysis is an investigation of the cointegration properties of the variables m_t , y_t , r_t , r_t^o , r_t^D , π_t , e_t^{US} , e_t^D . That means we investigate whether linear combinations of all or subsets of these nonstationary variables exist which are stationary. Such stationary linear combinations can very often be interpreted as long-run economic relations. Here we are especially interested in finding a long-run money demand relation.

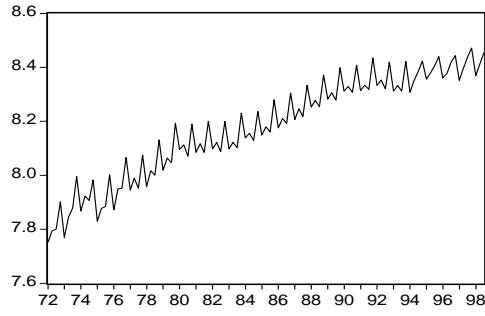
We applied JOHANSEN'S (1995) maximum likelihood approach for the whole period. Due to the presumed break in m_t around 1983, we include the step dummy $S83q1_t$, which is one starting with 1983(1) and zero before. The test is performed within a vector error correction model of order one, which was chosen following the Schwarz information criterion, allowing for a linear deterministic trend and including centered seasonal dummies. Since dummy variables as $S83q1_t$ influence the asymptotic distribution of the likelihood-ratio test statistic for the cointegration rank we have used the program DISCO (see JOHANSEN AND NIELSEN, 1993) to simulate critical values for this particular case. The likelihood-ratio test indicates that there exist three cointegrating relations, which would be consistent with the results found by GENNARI (1999), RINALDI AND TEDESCHI (1996) and JUSELIUS (1998). Using a 5% significance-level, the hypothesis of a cointegration rank of zero, one and two can be clearly rejected, while a rank of three cannot be rejected, even if the usual critical values are considered.

⁶All computations are done with EViews 3.1.

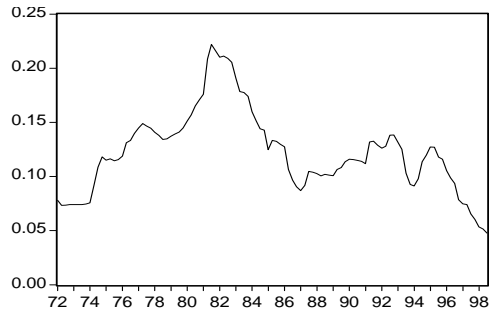
Figure 1: The time series used



m_t



y_t



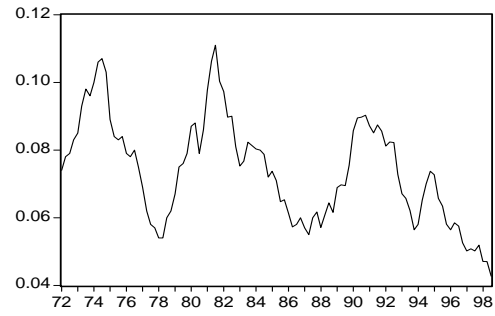
r_t



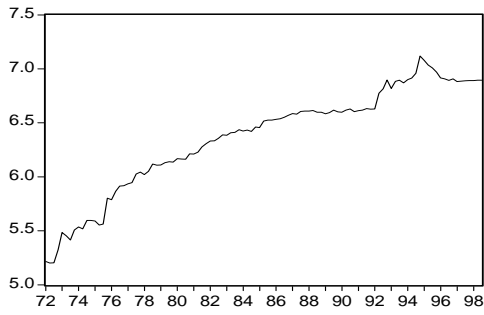
r_t^o



π_t



r_t^D



e_t^D



e_t^{USA}

3 Specification and Estimation of a Money Demand Function

Due to all the structural changes in the Italian monetary system we do not try to model a money demand equation within this highdimensional vector error correction model, but use instead a single equation approach to gain more flexibility for specifying the various changes which took place during the period under observation. Moreover, in monetary policy very often single equations are preferred, because they are easier to understand and interpret.

We use the ENGLE-GRANGER (1987) approach to estimate a long-run money demand function. Due to the structural changes around 1983, especially the switch from CTI to M2 as intermediate target we include the shift dummy $S83q1_t$ described above. The Lira/DM exchange rate is included only after 1992, when Italy left the ERM and the exchange rate was not restricted to stay within the given band. Thus, e_t^D is multiplied by the shift dummy $S92q1_t$, which is zero before 1992(1) and one after. Finally, this estimation approach leads to the following specification (standard errors are given in parantheses beneath the estimated coefficients)

$$\begin{aligned}
 m_t = & 4.35 + 0.62y_t - 1.04r_t + 5.84(S83q1_t \cdot r_t^o) - 2.17(S83q1_t \cdot \pi_t) \\
 & \quad \quad \quad (0.51) \quad (0.07) \quad (0.20) \quad (0.87) \quad (0.48) \\
 & \quad \quad \quad - 0.83r_t^D - 0.020(S92q1_t \cdot e_t^D) - 0.32S83q1_t + \hat{z}_t \quad (1) \\
 & \quad \quad \quad (0.32) \quad (0.002) \quad (0.06) \\
 \bar{R}^2 = & 0.60 \quad \quad \quad SE = 0.039 \quad \quad \quad JB = 0.25(0.88) \\
 T = & 107 \quad [1972(1) - 1998(3)].
 \end{aligned}$$

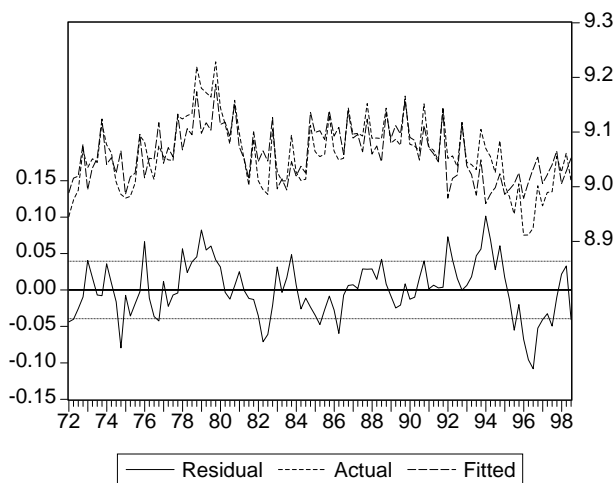
Since all variables are I(1) the usual t -distribution cannot be applied. The standard errors should be used as a more descriptive device. For the Jarque-Bera (JB)-test the p -value is given in parantheses. This statistic is computed to test if the residuals are normally distributed. Looking at the p -value of 0.88 the null hypothesis of normally distributed residuals cannot be rejected.

The residuals \hat{z}_t of regression (1) were tested for cointegration by applying the augmented Dickey-Fuller unit root test. The null hypothesis of no

cointegration implies that \hat{z}_t is I(1), whereas the alternative hypothesis of cointegration implies that \hat{z}_t is stationary. The resulting t -statistic amounts to -5.19. The critical values, which are tabulated by MACKINNON (1991) only include the case of five nonstationary regressors, whereas here we have to deal with six I(1)-variables as regressors. Taking the critical value for five regressors, which is -4.70 (5%-level) and -5.24 (1%-level) respectively, as an approximate value, it can be concluded that \hat{z}_t is stationary and that therefore equation (1) forms a cointegration relation.

The actual and fitted values as well as the residuals \hat{z}_t are depicted in *Figure 2*. It seems that the break in 1983 is very well captured by our specification. But it also can be seen that there are some turbulences at the end of the sample, around 1995/1996.

Figure 2: Residuals of the static regression (1)



All coefficients in the long-run money demand equation (1) show the signs which are postulated by economic theory. The income elasticity of 0.62 is clearly smaller than one and more in the neighbourhood of an income elasticity of 0.5, predicted by the BAUMOL-TOBIN framework. The coefficients of r_t , r_t^o , π_t and r_t^D , are semielasticities. Assuming a long-term Italian rate of 12% leads to an interest elasticity of -0.12 ($= -1.04 \cdot 0.12$). Assuming 5% rates for r_t^o , π_t and r_t^D leads to elasticities of 0.29, -0.11 and -0.04 respectively. Moreover, important variables in a money demand equation as the

own rate of money and the inflation rate are only effective since 1983 after the 'divorce' and the use of M2 as intermediate target. The results further show that currency substitution, modelled by the German long-term interest rate plays an important role during the whole period. After 1992(1), when the Italian Lira left the ERM we also found a significant influence of the Lira/DM exchange rate, whereas the Lira/US Dollar exchange rate does not contribute significantly to the performance of the regression and has therefore not been included. If we would exclude $S92q1_t \cdot e_t^D$ from equation (1), we would not be able to model the break in M2 in 1995 and 1996 in a satisfactory manner. Thus, currency substitution with regard to Germany is even stronger since Italy left the ERM.

The specification of an error correction model (ecm) is the next step in our analysis. Following the Engle-Granger procedure the lagged residuals \hat{z}_{t-1} resulting from the static regression (1) are used as error correction term within the error correction equation. We allow for up to four lags for the changes of the variables and include $\Delta S83q1_t$, $I74q4_t$ (which is one for 1974(4) and zero elsewhere) as well as seasonal dummies d_{it} , $i = 1, 2, 3$. $I74q4_t$ is used according to JUSELIUS (1998) since there was a sharp decline in the money stock and real GDP induced by very tight fiscal policy and the introduction of quantitative restrictions on bank lending. Successively eliminating the variables with the lowest t -values results in the following specification (t -statistics are given in parantheses beneath the estimated coefficients)

$$\begin{aligned} \Delta m_t = & -0.21 \hat{z}_{t-1} + 0.16 \Delta m_{t-1} + 0.35 \Delta m_{t-4} - 0.25 \Delta y_{t-1} - 0.18 \Delta y_{t-2} \\ & \quad \quad \quad (-4.57) \quad \quad \quad (2.05) \quad \quad \quad (4.82) \quad \quad \quad (-3.15) \quad \quad \quad (-2.43) \\ & - 0.18 \Delta y_{t-4} - 0.25 \Delta(S83q1_t \cdot \pi_t) \\ & \quad \quad \quad (-2.26) \quad \quad \quad (-2.17) \\ & + 0.074 - 0.12 d_{1t} - 0.068 d_{2t} - 0.095 d_{3t} - 0.090 I74q4_t + \hat{u}_t \quad (2) \\ & \quad \quad \quad (6.67) \quad \quad \quad (-5.46) \quad \quad \quad (-5.14) \quad \quad \quad (-6.19) \quad \quad \quad (-5.14) \end{aligned}$$

$$\begin{aligned} \bar{R}^2 = 0.89 & \quad \quad \quad \text{SE} = 0.017 & \quad \quad \quad \text{JB} = 0.67 (0.71) \\ \text{LM}(1) = 1.25 (0.27) & \quad \quad \quad \text{LM}(4) = 0.55 (0.70) & \quad \quad \quad \text{LM}(8) = 0.60 (0.78) \\ \text{ARCH}(1) = 1.64 (0.20) & \quad \quad \quad \text{ARCH}(4) = 0.99 (0.41) & \quad \quad \quad \text{RESET}(1) = 7.74 (0.01) \\ & \quad \quad \quad \text{T} = 106 [1972(2)-1998(3)] \end{aligned}$$

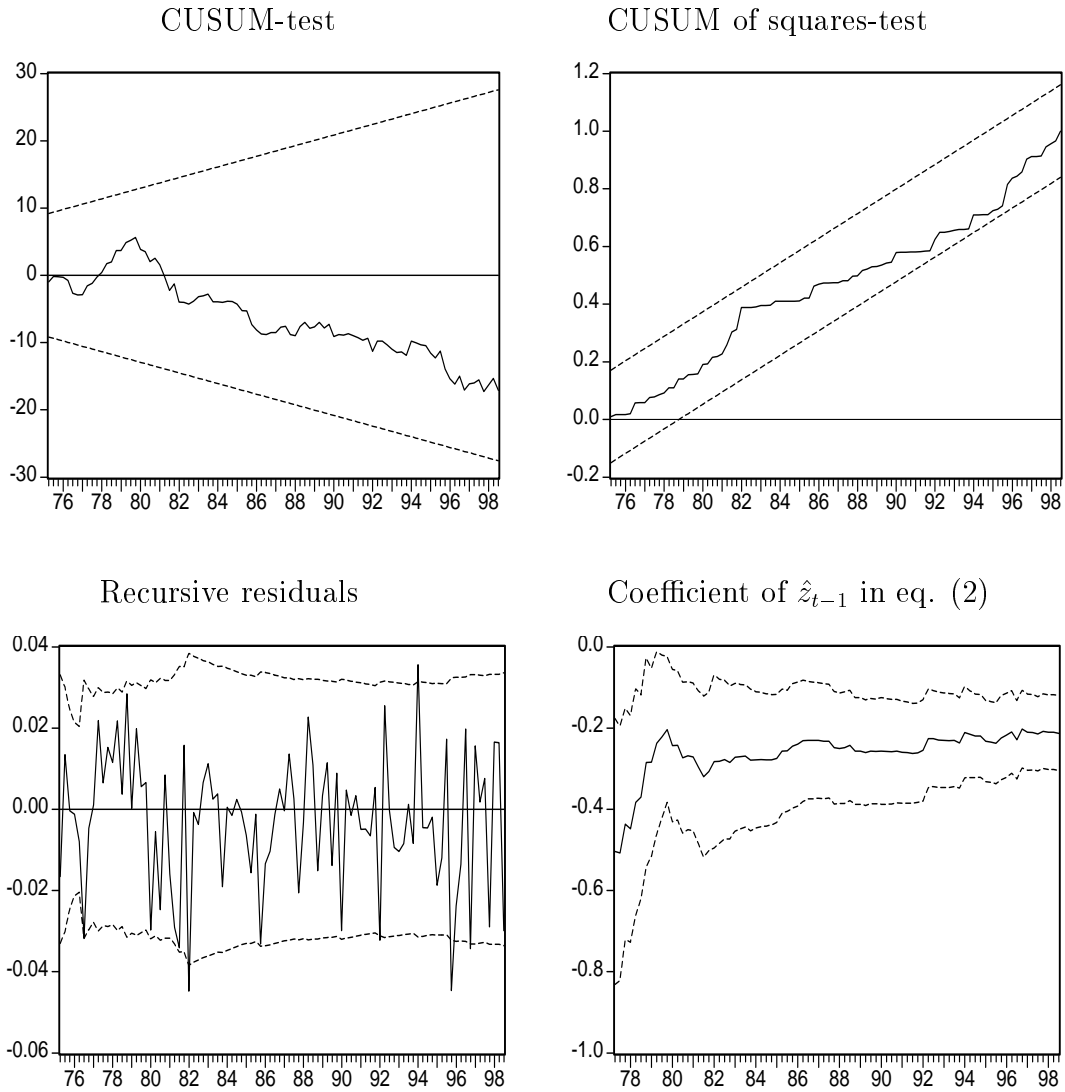
For the test statistics the p -values are given in parantheses. The diagnostic tests indicate that the model is well specified. The assumption of normally distributed residuals cannot be rejected (JB). Lagrange multiplier (LM) tests of no error autocorrelation against autocorrelation of at most order $k = 1, 4, 8$ do not indicate any problems concerning autocorrelated residuals. The model also passes the LM tests against conditional heteroskedasticity (ARCH(k)) of at most order $k = 1, 4$. Only the RESET(1)-test leads to the assumption that there may be some kind of nonlinearity. The t -statistic of the error correction term amounts to -4.57, which can again ensure the existence of a cointegration relationship. Moreover, the negative sign of the coefficient shows that the adjustment process leads into the right direction, that means excess money is followed in the next period by a reduction of the money stock.

The stability of the estimated equation has been checked using the CUSUM-test, the CUSUM of squares-test, the plot of the recursive residuals, and the plot of the recursive estimates of the coefficient of the error correction term \hat{z}_{t-1} in equation (2).⁷ All these results are shown in *Figure 3*. (For the CUSUM- and CUSUM of squares-test the 5% critical lines, for the recursive residuals and estimates the two-standard-error bands are shown.)

The CUSUM- and the CUSUM of squares test do not indicate any instability of the estimated equation. By examining the plot of the recursive residuals only three values which lie not very far outside the two-standard-error bands can be noticed. The recursive estimate of the coefficient of the error correction term in equation (2) also shows no significant instability. Thus a stable money demand equation seems to exist in Italy for the period 1972 to 1998.

⁷Since $\Delta(S83q1_t \cdot \pi_t)$ is zero before 1983(1) we generated a white noise process with zero mean and a variance of 0.00001 for 1972(2)-1982(4) to be numerically able to compute the recursive estimates from the beginning.

Figure 3



4 Conclusions

The period from 1972 to 1998, here under observation, is characterized by major institutional changes in the design of Italian monetary policy and in the financial system. Despite this fact we are able to detect a stable money demand function for the whole period by using the Engle-Granger approach. We estimate the long-run relation as well as the short-run dynamics through

an error correction model. What has been evident is the influence of currency substitution, especially between Italy and Germany, which has been taken into account by the Lira/DM exchange rate and a German interest rate. Especially the Lira/DM exchange rate enables us to model the structural break in 1995 and 1996 in M2. Other studies concerning the Italian money demand do not consider currency substitution and do not include the whole sample and especially the recent data. One corollary of the relevance of currency substitution found by us is that Italian monetary policy may have exacerbated the depreciation of the currency in 1994-1995, when the money stock was not growing (or growing very little) and money demand was falling by more.

The estimated magnitude of the income elasticity (0.62) is comparable to the results found by ANGELINI ET AL. (1994) or BAGLIANO (1996). ANGELINI ET AL. (1994) estimate an elasticity to domestic demand of around 0.6, depending on the frequency of the data and an elasticity of the own rate of money that is higher than the one of the alternative yield, a result which also holds in our case. The income elasticity found by BAGLIANO (1996) amounts to about 0.8, depending on the different estimation techniques. These empirical results are also consistent with the BAUMOL-TOBIN framework or with the development of close substitutes to M2 in the Italian financial system.

It seems that in this period, with which we are concerned, the monetary aggregate M2 fulfills the necessary conditions to be used as intermediate target for Italian monetary policy.

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