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ELENA GENNARI

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ELENA GENNARI

BADIA FIESOLANA, SAN DOMENICO (FI)

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European University Institute
Badia Fiesolana
I – 50016 San Domenico (FI)
Italy

Estimating Money Demand in Italy 1970 -1994

Elena Gennari *
European University Institute
Department of Economics

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Abstract

A money demand relationship for Italy is estimated from 1970 to 1994 within a cointegrated VAR framework. Changes in the money market due to an important financial innovation process are introduced in the cointegration space through a Logistic Smooth Transition function tested and estimated at an earlier stage using Engle-Granger cointegration analysis. Results suggest the importance of such a non linearity to acheive a better identification of the long-run equilibria although differences in variability between the periods pre and post-1983 still emerge.

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1 Introduction

The purpose of this paper is to estimate the demand for money in Italy from the beginning of the 70's to the middle of the 90's. These years, in fact, have been characterized by important changes in the Italian money market structure: progressive liberalization of capital movements, the 1975 reform of the Treasury bill market and the introduction of new financial instruments. The characteristics of the market and the role of money within the system have been altered by these changes. In fact, from a dual role of medium of exchange in transactions, and an alternative riskless asset for investment, money is now essentially used only for the first purpose, since there are many others alternative zero-risk financial assets offering a higher yield.

In particular, the present work is focused on the role of the financial innovation in money demand and on the importance of modelling it within the system to achieve a better determination of the long-run equilibria. Money demand is, in fact, estimated within a system and identified as one of the long-run relationships linking these variables. The empirical analysis is done using the concepts of cointegration analysis both in univariate (see Engle and Granger, 1987) and in multivariate contexts (see Johansen 1995). Nonlinear modelling is also introduced in estimating a logistic smooth transition function (see Teräsvirta, 1996) to account for financial innovation in the money market. The results confirm the importance of such a nonlinearity to achieve a better determination of the money demand equation.

The paper is organized as follows. Section 2 summarizes previous works in the field. Section 3 concerns basic economic theory behind the estimation of a money demand relationship while Section 4 explains very briefly the econometric methodology employed to estimate the system of variables. Section 5 presents the set of data used. Section 6 gives an account of the role of financial innovation for the Italian money market, describes how this phenomenon has been dealt with in the literature and suggests a possible way of modelling it. In Section 7 we estimate a logistic smooth transition function using univariate cointegration analysis while in Section 8 we perform Johansen cointegration analysis on the system with the estimated logistic trend. Finally, in Section 9 we present the final model obtained from the system and in Section 10 we conclude.

2 Previous work

The evidence found of the instability of the estimated money demand functions for various countries during the seventies has generated a lot of empirical research work in the field. To focus only on the most recent, in the Italian case, following the work of Cuthbertson and Taylor (1987), Bagliano and Favero (1992) estimate a forward looking model from 1964 to 1985 which is able to solve the seventies episode of instability in money demand but fails to account for other important instability phenomena.

More recently, an important contribution is the work of Angelini *et al.* (1994) who estimate a single equation model for quarterly data from 1975 to 1991 and for monthly data from 1983 to 1991. They try to deal with the instability of the estimated money demand equation for quarterly data adjusting the interest rates in order to account for the learning effect, risk and the decreasing illiquidity premium but none of these devices solves the problem. They finally allow for two different scale variables within two subsamples. In the period 1975-79 the net financial asset is the variable with a significant effect on money demand while from 1983 to 1991 this is substituted by domestic demand. The shift from one regime to the other one is captured by a weighted average where coefficients give progressively more weight to the second scale variable, emphasizing the transaction motive and reducing the speculative one.

Within the multivariate framework, a recent contribution is that of Rinaldi and Tedeschi (1996) who estimate a VAR with monthly data over the period 1983-1991 and, using Johansen's cointegration procedure, identify the money demand relationship as one of the vectors of the cointegration space. In particular, the work of Rinaldi and Tedeschi (1996) aims to check if the univariate analysis of Angelini *et al.* (1994), which conditions on all the other variables in the system, is legitimate or not.

Building up a VAR(2) and identifying three cointegration relationships, they reduce the system to a simultaneous equation model finding evidence of long-run exogeneity of inflation, income, the yield on M2 and the yield on Treasury Bills with respect to the parameters of the money equation.

Finally, one of the last works in the field is that of Bagliano (1996) who estimates a money demand function using monthly seasonally adjusted data from 1983 to 1991. Johansen's cointegration procedure is performed on a system of four variables: real money (M2 definition), total final expenditures, the after tax yield on Treasury bills averaged over three, six and twelve month maturities and the after tax own return on M2. The inflation rate (which is represented by the first difference of the logarithm of the consumer price index) is excluded from the long-run relationships and introduced as exogenous in the explanation of the short run dynamics[1]. Two cointegration vectors are found amongst the four I(1) variables, one linking real money to total final expenditures, which is interpreted as a money demand relationship, and the other linking the two yields while price homogeneity (i.e. $(m - p)$) and exogeneity of the inflation rate are successfully tested.

With respect to these previous works, the present estimation of the money demand considers a longer sample which includes an important transition period of the money market which is directly included in the long-run equilibria of a system of variables. Results can be compared with the quarterly model of money demand estimated by Angelini *et al.* (1994) although weak exogeneity of the inflation rate is rejected in the present context, so that the system cannot be opened and reduced to a single-equation model.

3 Economic Theory

Generally speaking, money has two basic roles within an economy: it acts as a medium of exchange in transactions and as a zero-risk asset in

[1] The choice is justified on the basis of the stationary behaviour of the series.

portfolio choices [2].

The demand for money as a pure transaction medium depends basically on the price level and on a measure of expenditure, which is normally given by income. The demand for money as a financial asset depends on the rate of return on the money itself and on the yields of alternative investments, like the yield on bonds, or, if we consider real assets amongst the possibilities, on the inflation rate. It obviously depends also on the wealth to be allocated in the portfolio.

A long-run relationship for money demand will thus include the price level, income as a scale variable, the own rate of return on money, the yields on alternative assets and could take different forms. If we are focusing on a broad monetary aggregate then both roles of money are present and the relationship could be of the kind:

$$(m - p) = \beta_1 + \beta_2 y + \beta_3 i_m + \beta_4 i_b + \beta_5 \Delta p \tag{1}$$

where m is the logarithm of nominal money, p is the logarithm of the price level[3], y is the logarithm of income, i_m is the own rate of return on money, i_b is the yield on bonds and Δp is the inflation rate. Typically, the equation reported above will have $\beta_2 > 0, \beta_3 > 0, \beta_4 < 0, \beta_5 < 0$ (see, for example, Ericsson and Sharma, 1996). A coefficient $\beta_2 = 1$ would be consistent with the quantity theory of money while $\beta_5 = 0$ will exclude any role for inflation as a determinant of the money demand; $\beta_3 = -\beta_4$ will, finally, imply dependence on the pure opportunity cost. These very general *a priori* considerations about the determinants of the demand for money constitute the general guidelines for the choice of variables to be included in the system and for the identification of the money demand relationship in the estimated VAR when applying the cointegration analysis.

[2]For a survey of the literature on money demand see, for example, Goldfeld and Sichel (1990).

[3]Instead of considering real money we may also test the validity of the homogeneity condition. In the present work we will assume the restriciton to hold. However, in a different work (see Juselius and Gennari, 1998), evidence suggests that this condition does not hold for the full sample but only if we restrict estimation to the second part of the period.

4 Testing long-run relationships through cointegration

The modelling of I(1) time series that share long-run common stationary equilibria has been traditionally done using either single-equation models or Vector Autoregressive models (VARs) (see, *inter alia*, Banerjee *et al.*, 1994). In the first case, a simple linear regression is fitted to the levels of the variables and the residuals are then tested for stationarity to assess the existence of cointegration amongst the set of variables (see Engle and Granger, 1987). The OLS estimation delivers estimators for the parameters which are superconsistent, given the I(1) property of the series, but with non-normal limiting distributions. However, this procedure does not allow for more than one relationship linking the variables. This may, in fact, be a linear combination of more than one cointegrating vector when such exist.

To overcome these problems, a more comprehensive econometric procedure based on the estimation of VARs has been developed to disentangle the effects of the various long-run equilibria. These models characterize the joint behaviour of a group of variables conditional on their past values and, possibly, on a group of deterministic variables which may include the constant term, the linear trend, seasonal dummies and event-specific dummies (like impulse or step dummies).

A VAR model with p variables takes the following form:

$$X_t = \Pi_1 X_{t-1} + \ldots + \Pi_k X_{t-k} + \Phi D_t + \epsilon_t \qquad t = 1, \ldots, T$$

which has fixed values $X_{-k+1}, \ldots, X_0$ and where ϵ_t is a p dimensional normal process $N_p(0, \Omega)$.

If the I(1) variables that we are modelling have r stationary equilibrium relationships amongst them[4], then we can write the system in the stationary Vector Equilibrium Correction form (VE_qCM) (Johansen,

[4]The variables that have this property are said to be 'cointegrated'.

1988) which will be given by:

$$\Delta X_t = \Pi X_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \Phi D_t + \epsilon_t$$

where $\Pi = \sum_{i=1}^{k} \Pi_i - I$ and $\Gamma_i = -\sum_{j=i+1}^{k}$. The matrix Π has reduced rank equal to r and can be decomposed as follows:

$$\Pi = \alpha\beta'$$

where α and β are $p \times r$ full rank matrices. This gives rise, in the VE_qCM form, to the term $\alpha\beta' X_{t-1}$, i.e. the α matrix multiplied by $\beta' X_{t-1}$ which is a vector of r I(0) linear combinations of the variables in the system.

Taking into account this decomposition we have:

$$\Delta X_t = \alpha(\beta' X_{t-1}) + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \Phi D_t + \epsilon_t \tag{2}$$

(see Banerjee *et.al.*, 1994 and Johansen, 1995).

In order to test for the number of cointegration relationships amongst the variables, *i.e.* for the rank of the matrix Π, we can use two tests developed by Johansen (1988, 1991), one based on the sum of the first $(p-r)$ eigenvalues (trace statistic) and the other based on the $(p-r)$-th eigenvalue (maximum eigenvalue statistic)[5]. These respectively test that the first $(p-r)$ eigenvalues are equal to zero and that the $(p-r)$-th eigenvalue is equal to zero starting from $r = 0$ and increasing r.

Once the number of relationships r has been determined, the coefficients of these will be given by the eigenvectors associated to the first r eigenvalues, and restrictions can be imposed on the coefficients to test alternative *a priori* theory-based hypothesis on the long run behaviour of the variables.

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[5]These eigenvalues solve the problem of maximisation of the likelihood in the reduced rank regression arising from model (2).

5 The Data

The data used in the estimation of this model of money demand are given by the logarithm of real money supply, M2 definition, calculated as the difference of the logs of the nominal money supply and the consumer price index; the quarterly rate of inflation, which is derived as the first difference of the logarithm of the consumer price index; the own rate of return on M2; the bond yield and the final domestic demand as scale variable. The own rate of return on M2 and the bond yield, for comparability with the inflation rate, have been divided by 100, to express them in absolute terms, and by four to make them quarterly rates of return.

The series of M2 money supply, the bond yield (government bond yield) and the inflation rate (rate of change of CPI-index) are OECD series while the final domestic demand is released by CENT-ISTAT[6].

All series are quarterly, seasonally unadjusted and the estimation sample extends from 1970(2) to 1994(4).

The graph of the series is reported in figure 1 together with money velocity. While final domestic demand has quite a regular seasonal pattern, real money supply shows two different seasonal behaviours before and after approximately 1981. This irregularity in seasonality that characterizes the 70's is probably linked to the situation in the exchange rate market following the two oil shocks.

Inflation is very low at the end of the 60's but the strikes of 1969, with the consequent signing of the new national labour contract, and the first oil crisis of 1973 push up once more the price growth. As we can see all the 70's are characterized by rising inflation which, after the second oil shock in 1979, reaches a yearly rate of 22%. At the beginning of the 80's inflation starts falling and in 1987 its level is just above 4%.

Both the return on money and the bond yield increases until 1981-82 and then starts falling while money velocity undergoes a substantial shift in mean which starts around 1979 and ends around 1982.

[6]The own rate of return on M2 has been kindly provided by R. Rinaldi of Bank of Italy, coauthor of a previous work on money demand.

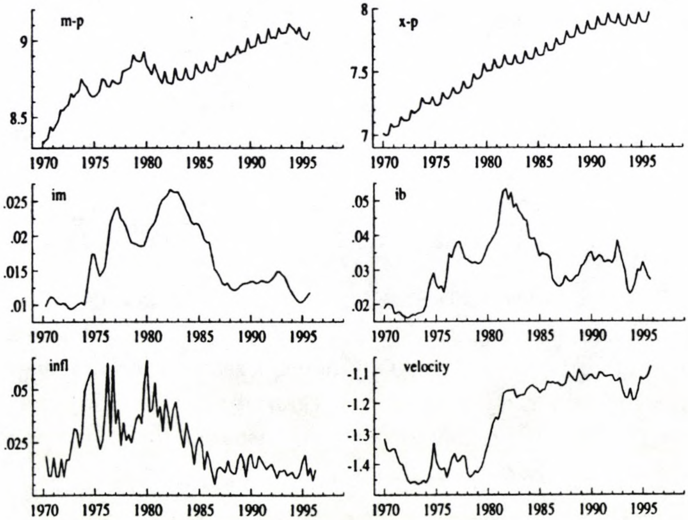

Figure 1: The series

6 The role of financial innovation

In the last 25 years the characteristics of the Italian money market have changed completely. The money market has, in fact, undergone a substantial process of financial innovation in two different aspects: a reform of the money market and the introduction of new financial instruments, and a process of modernization of the market. In the first category, the creation of a large market for government securities starts in 1975 with the Reform of the Treasury Bills Market. This reform concerns, in particular, the auctions where the Bank of Italy is obliged to participate as a residual buyer together with the other agents. The reform introduces also a floor price for every auction whose setting is left to the Treasury.

Together with the already existing BTP (Treasury bonds) and BOT (Treasury bills), in 1977 another type of government security is introduced, the Treasury's credit certificates with floating rate (CCT), whose indexation mechanism was revised in 1981, and in 1988 fixed rate Treasury option certificates (CTO).

The following years are characterized by the introduction of competitive bid auctions for three, four and twelve month Treasury Bills and, from 1988 onward, by the abolition of the floor price on the auctions. This date signals also the starting point of debt management aimed at lengthening the average maturity of the debt through the first issue of 5 and 7-year CCT[7]. It was followed in 1990 and in 1991 by the first issue of 7 and 10-year BTP and in 1993 by the first issue of 30-year BTP.

As far as the modernization of the market is concerned, an important step towards the creation of more developed market for government securities takes place in 1988: the introduction of the screen-based secondary market (for government securities). From this point on the money market underwent a process of innovation and modernization: the reform of centralized securities accounts (CAT) held with the Bank of Italy by

[7]In 1987 the new upward trend in inflation caused an increase in the demand for short term securities and a fall in subscribtion of long-term bonds so that the average maturity of government securities was 8 months at the end of 1987, see Passacantando, 1996.

banks, security firms, etc...(1990); the launch of BTP and Eurolira futures on LIFFE (1991-1992); the creation of the Italian Futures Market (MIF) for BTP futures (1992); the launch of options on BTP futures (1994) and, finally, in July 1994, the Reform of the screen based market (MTS) where Treasury securities are traded on the stock exchange screen-based system (see Passacantando, 1996, App.).

But the most relevant event that changed the money market during this period was, probably, the so-called 'divorce' between the Bank of Italy and the Treasury in July 1981. According to this agreement, the Bank of Italy was no longer obliged to act as a residual buyer in the auctions of the Treasury bills and the overdraft on the account of the Treasury with the Central Bank was set at a maximum of 14% of the total anticipated expenditures. However, the change was not abrupt: the Central Bank continued to guarantee a support to the Treasury until approximately 1983 (see Passacantando, 1996, p.90).

The effects of the underlying financial innovation process can be perceived by looking, in particular, at the upward shift of money velocity (fig.1) from 70's to 80's . This upward movement, shown by money velocity, is clearly not abrupt and could be represented by a logistic-type trend that highlights the smooth transition from one regime to the other: as a matter of fact, agents take some time to learn about the new instruments available on the market.

This kind of trend has, in fact, been used many times in the literature to approximate the learning process associated to financial innovations (see, for instances, Baba *et al.*(1992), Hendry and Ericsson (1990), Muscatelli and Papi (1990) and Vaciago and Verga (1989)). Hester (1981) says that '...innovations probably tend to alter observed relations between macroeconomic variables in [a] highly nonlinear manner...The diffusion of an innovation through an industry might reasonably be approximated by a logistic function that applies to the slope of some behavioural relation.'. Hence, the class of the Logistic Smooth Transition functions (LSTR) can be fruitfully employed to introduce modelling of the effects of market changes into the underlying long-run relationships.

In the money demand equation, the S-shaped type behaviour can be present in various coefficients, although what changes in general is either the reaction of the agents to the opportunity cost of holding money or the level of investment in the new instruments.

If we let (1) represent the money demand relationship, as described before, what is generally thought to change is either the constant term β_1 or the coefficients $\beta_3, \beta_4, \beta_5$ (or both). In the first case, agents are supposed to have always the same reaction to the yields on the various assets but the lack of profitable and riskless financial instruments does not push them to move capital from money in the broad sense to another form of investment. In the second case, on the other hand, agents have a different reaction to the spread between the yield on the other assets and the yield on money, a reaction which changes as new instruments are available on the market. Examples of both interpretations can be found in the literature. Vaciago and Verga (1989), for example, introduce financial innovation through a separate logistic trend while Baba *et al.* (1992) multiply this by the yield on M2 to obtain a learning-adjusted yield.

The effects of the financial innovation process have been represented (see Angelini *et al.* (1994)) also as a shift in the role of money: from money used both in transactions and as an alternative financial investment, to a pure means of payment role.

7 Testing and estimating the LSTR function in the cointegration space

Logistic smooth transition functions represent a class of nonlinearities used to model transitions from one regime to another which are not discrete but smooth. This kind of nonlinearity, as proposed by Maddala (1977), are represented by a logistic function of a transition variable s_t which can be a function of other variables or simply a function of time. The logistic function depends, furthermore, on a location parameter c and a slope parameter γ that describes, respectively, when the function changes concavity and how rapid the transition is.

The LSTR function we are going to use in the present paper[8] is represented by the following monotonically increasing function of time:

$$G(t, \gamma, c) = \{1 + \exp[-\gamma(t - c)]\}^{-1}$$

where $\gamma > 0$ is imposed for identification.

For a linear model $y_t = \theta' x_t + u_t$, the corresponding nonlinear model takes the following form:

$$y_t = \varphi' x_t + \theta G(\ t, \gamma, c) + w_t$$

which means that we can have a linear part, represented by $\varphi' x_t$ and a nonlinear part given by $\theta' G(\ t, \gamma, c)$. Notice that, in our case, the constant is the only term multiplied by the logistic since, as shown below, an approximation to this function has pointed out that this nonlinearity is sufficient to achieve better stationary properties of the cointegration space.

[8]This is not the only LSTR function we can use. For a full description of the whole class of functions, see Teräsvirta (1996).

Testing

The econometric theory for testing and estimating nonlinear models amongst I(1) variables is not fully developed. Given the simple type of nonlinearity used in the present context, which is a deterministic function of time, we will follow a more descriptive approach.

In order to test for the presence of such a nonlinear logistic trend in the cointegration space and then estimate it, we use the Engle-Granger framework described above specifying the following nonlinear univariate model in levels:

$$(m-p)_t = \beta_1+\beta_2 y_t+\beta_3 i_m+\beta_4 i_b+\beta_5 \Delta p+\beta_6\{1+\exp[-\gamma(t-c)]\}^{-1}+w_t^* \quad (3)$$

The parameters γ and c are identified only under the alternative hypothesis of the presence of the nonlinearity (i.e. $\beta_6 \neq 0$) but not under the null. Therefore we have tested for nonlinearity using a third order Taylor series approximation to the smooth transition function, i.e. test a model with the terms t, t^2, t^3.[9] As we are dealing with a regression amongst non stationary variables, the standard asymptotic results are not valid and thus the decision as to whether nonlinearity is needed is taken looking at the stationary properties of the residuals, i.e. accept nonlinearity if the addition of the powers of t helps achieving more stationarity in the residuals.

The critical values of the DF and ADF tests depend on the presence of deterministic terms like the ones we introduce thus we decide whether the hypothesis of stationary can be accepted by looking at the graphs and correlogram of the residuals. The results seem to point to the importance of this form of nonlinearity to achieve residuals which seem more stationary. The logistic trend is producing, in fact, more or less the same result as a step dummy for a change in regimes. Without the intervention of this deterministic term there is less evidence of cointegration (using Engle-Granger procedure) because of the shift which makes

[9]If only the term t^2 were used, a different LSTR model would be tested, i.e. a model of this form:
$G(t, \gamma, c) = \{1 + \exp[-\gamma(t - c)^2]\}^{-1}$

the equilibrium appear less stationary. In other words, we assume that there is, at least, one cointegration vector amongst the series (i.e. there is cointegration) but a nonlinear deterministic term enters the long-run equilibria. In fact, from the graphs reported below (fig.2), we can see that introducing the term t^3 caused the residuals to change their behaviour completely and the correlogram decays faster. For sake of comparison we have also graphed the residuals of the estimation using the logistic function illustrated in the following paragraph. This is done to show that an LSTR function is actually the nonlinearity which delivers the most stationary residuals.

Estimation

Given these results we then estimate the logistic function with the use of nonlinear Least Squares applied to model (3). The estimation process delivers the following LSTR function[10]:

$$G(t,\widehat{\gamma},\widehat{c}) = \{1+\exp[-\underset{(0.030072)}{0.17230}\,(t-\underset{(1.6756)}{47.717})]\}^{-1}$$

8 The long-run structure

8.1 The VAR

Once the logistic trend has been estimated with nonlinear least squares, we introduce it in the VAR[11], estimate the system and use Johansen cointegration procedure to determine the rank r and to possibly identify a long-run money demand amongst the cointegration vectors.

Given that the data are quarterly, we start with the estimation of a VAR with five lags for the full sample and, as real money is affected by seasonality, we introduce a set of centred seasonal dummies which are

[10]Estimates are obtained using PcFiml, PcGive (Doornik and Hendry, 1995).

[11]The variable which corresponds to the new trend is called 't_L' and it is restricted to lie within the cointegration space.

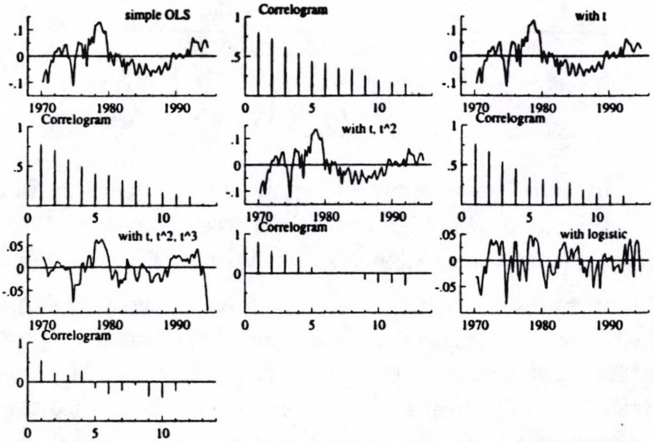

Figure 2: OLS residuals and correlograms

orthogonal to the constant term[12]. Furthermore, the estimates of the unrestricted VAR include also two impulse dummies: one for 1986q2, introduced mainly to capture some instability in the equations for the yield on bonds, which corresponds to a drastic reduction of the discount rate (3% in less than three months) operated in Spring 1986; the second one for 1974q4, which is related to some instability in the bond market due to the consequences of the first oil shock (see Caranza and Cottarelli, 1987).

The diagnostics, in the form of single equation statistics, that we

[12]In fact, in the moving-average representation of the process, a deterministic variable would give rise to the following term:

$$C\sum_{i=1}^{t}\Phi D_i + C(L)\Phi D_t$$

If D_t is a not centred seasonal dummy, the first term will be a linear trend, while the second will give a seasonally varying mean (see Johansen, 1995, p.84).

Table 1: Full system statistics

statistic	value	p-value
Vector portmanteau 10 lags	196.73	
Vector AR 1-5 F(125,132)	1.1500	[0.1991]
Vector normality $\chi^2(10)$	5.6378	[0.8447]

do not report here and of vector statistics, that we report in Table 1[13], indicate a well specified model, apart from some autocorrelation in the equation for domestic demand which is significant, however, only at 5%.

The presence of the logistic trend within the cointegration space alters the critical values so that the maximum eigenvalue and the trace tests are not reliable in this case. We have calculated, with the help of the program Disco (Johansen and Nielsen, 1993), the critical values for a linear approximation to the logistic (i.e. a broken trend) in the case of the trace test. In Table 2 we then report only the values of this statistic with the ** and * corresponding to significance at 5% and 1% critical level.

The evidence from the cointegration analysis is not always so clear-cut. In this case we rely also on the size of the eigenvalues, on the roots of the companion form matrix and on the graph of the cointegration vectors. The overall evidence seems to support the presence of three stationary equilibria and thus we test for meaningful restrictions on a set of three long-run relationships looking, in particular, for a money demand relationship.

8.2 Testing restrictions on β

Economic theory has an important role in this kind of analysis because, at this stage, it drives the search over the various possible structures

[13]The tests for the full system diagnostic are vector extensions of single equation tests (see Doornik and Hendry 1997) and are given by:

-Vector Portmanteau statistic (see Lüthkepohl, 1991);

-Vector autocorrelation test (see Godfrey, 1988);

-Vector Normality test (see Doornik and Hansen, 1994).

Table 2: Cointegration test

Eigenv.	$\lambda-$trace	$H_0 : r$	p-r
0.440135	114.6**	0	5
0.215113	60.09*	1	4
0.154595	37.32	2	3
0.123406	21.54*	3	2
0.0928113	9.156**	4	1

of the long-run equilibria helping to identify economically interpretable explanations of the phenomena. Seeking to identify completely the cointegration space, we have not only tested for the presence of the money demand relationship, which is our original purpose, but also the existence of other important links between the modeled variables (see Juselius, 1997). In particular, we have tested for the presence of a *central bank reaction rule*, which links positively the spread between the short term yield and the longer term yield to the differences of the inflation rate from a target π^*, i.e.:

$$E[i_{mt} - i_{bt} - \beta_0(\Delta p_t - \pi^*) - \mu] = 0$$

where μ is a constant and β_0 is positive. This kind of relationship is often found in the literature on money market since many central banks use the discount rate as an instrument to control monetary conditions (see Juselius, 1992).

Another relationship for which we have tested is an *aggregate income relationship* which should be an equilibrium amongst the scale variable, the inflation rate and one of the two yields, taking into account both an IS curve and a short-run Phillips curve. The following formulation considers both possibilities:

$$E[y_t - \beta_1 t - \beta_2 i_{bt} - \beta_3 \Delta p_t] = 0$$

The IS case would have $\beta_1 \geq 0$, $\beta_2 < 0, \beta_2 = -\beta_3$, while the short-run Phillips curve would typically have $\beta_2 = 0$ and $\beta_3 > 0$.

Furthermore, yields are theoretically linked by two fundamental relationships. The first one tells us that the short interest rate depends on expected inflation (Fisher's parity). Introducing Δp_t as a proxy for $E(\Delta p_{t+1})$ we test a relationship of this kind:

$$E[i_{mt} - \Delta p_t] = 0$$

The second relationship linking the yields is the expectations hypothesis which predicts that the longer yield is determined by the shorter yield:

$$E[i_{bt} - i_{mt}] = 0$$

All these relationships are tested using restrictions on the coefficient of the cointegration relationships so that in the expressions above the disturbance term is, actually, a stationary variable. Rejection of a particular hypothesis means that it is not possible to find a long run stationary relationship of that type amongst the group of the modeled I(1) variables.

The final structure (see also fig.3) which we found reasonably well supported by the data is the following one:

$$(m-p) - \underset{(0.028897)}{1.1521}(x-p) = \underset{(0.77228)}{5.1038}(i_m - i_b) - \underset{(0.018304)}{0.36235}\ t_L$$

$$(i_m - \Delta p) = \underset{(0.0028635)}{0.013819}\ t_L$$

$$(i_b - i_m) = \underset{(0.12366)}{0.76887}\ \Delta p + \underset{(0.0038736)}{0.020840}\ t_L$$

$$LR-test \qquad \chi^2(3) = 2.9942 \qquad p-value = 0.3925$$

The first relationship is a money demand equation in which real money depends on the scale variable, whose coefficient is, however, different from -1, on the opportunity cost and on the logistic trend (t_L)[14].

[14]The value of the coefficients of the money demand relationship is a bit lower with respect to what has been found, for instance, by Rinaldi and Tedeschi (1996) and, for Greek data, by Ericsson and Sharma (1998).

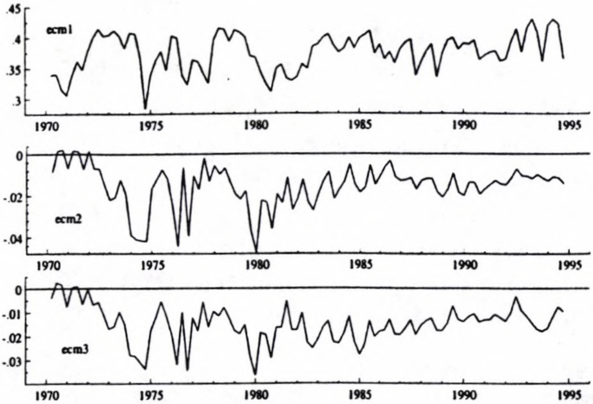

Figure 3: Cointegration relationships

The second relationship says that real yield on money is stationary around the logistic trend so that it is equal to the Fisher's parity expression with the addition of a term that accounts for a shift in the mean.

The third vector, whose restrictions are those of the expectation hypothesis, can be explained recalling that i_b is a medium term yield on government bonds. This means that the measure of inflation we are using here is a bad proxy for the expected inflation in the medium term and thus the spread on the two yields cannot be completely stationary. There is something more to be accounted for, i.e. increasing expectations on inflation which characterize the sample period and come out as an 'additional' inflation term in the third equation.

Finally, notice that, especially for the second and the third vector, there is a clear different variability pre and post 1982-83. To assess the significance of this difference, a formal likelihood ratio test is performed in par 8.4 splitting the sample at 1982q4.

8.3 Testing restrictions on α

Another important feature of the cointegration analysis is represented by the weights that these relationships have in the various equations. These are given by the elements of the matrix α. As we can see from the following table, real money seems to react only to disequilibria in the money demand relationship (equilibrium correcting) while inflation is pushed up by excess money supply. The third error correction term is only significant in the equation for the inflation rate and it basically says that when the opportunity cost is higher than what is compatible with the equilibrium level, inflation expectations are pushed up.

We can test various hypotheses on the parameters of the matrix α. A first interesting aspect is represented by the possibility of identifying long run weak exogeneity of a variable, or a group of variables, with respect to the parameters of the equilibrium relationships. If the three vectors do not have any influence on a particular variable, in which case all the weights will be equal to zero, then that variable is said to be long-run weakly exogenous for the long-run parameters and thus can be considered as driving the dynamics of the system as a whole. We have tested this particular hypothesis for final domestic demand and the bond yield (see Table 4), both without and with the imposition of the long run equilibrium structure, finding evidence of exogeneity for the first one but no strong support for the exogeneity of the second one. This means that the final domestic demand can definitely be considered a stochastic trend driving the system as the shocks to this variable 'cumulate in the system and give rise to the non-stationarity' (see Johansen, 1995, p.123).

This last observation introduces a second important aspect connected to testing hypotheses on the $\alpha's$: the interpretation of $\alpha'_{\perp} \sum_{i=1}^{t} \varepsilon_1$ as common trends driving the dynamics of the system. If a variable is, in fact, long run weakly exogenous then the matrix $\alpha_{\perp}$ will have one column that picks up only the cumulated residuals of a particular variable that will thus constitute alone one of the driving trends.

From the structure reported in Table 3, first imposing the long run exogeneity of $(x-p)$ and then deleting other insignificant coefficients, we

Table 3: Adjustment coefficients

	ECM_1	ECM_2	ECM_3
$\Delta(m-p)$	-0.325 0.096	-0.065 0.601	0.408 0.627
Δim	-0.006 0.003	-0.077 0.021	0.040 0.022
Δib	-0.027 0.015	-0.208 0.096	0.047 0.099
$\Delta infl$	0.114 0.046	0.376 0.288	0.671 0.300
Δxp	0.033 0.061	0.014 0.383	-0.497 0.400

Table 4: Exogeneity test

variable	LR-test	prob. value
$(x-p)$	2.2852 $\chi^2(3)$	0.5154
i_b	8.6123 $\chi^2(3)$	0.0349
$(x-p)$ i_b	4.6569 $\chi^2(6)$ 11.207 $\chi^2(6)$	0.5885 0.0822

can arrive at the following simplified matrix, whose structure is reflected also in the final estimation of the parsimonious system that we will derive in the following paragraph:

$$\begin{bmatrix} a_{11} & 0 & a_{13} \\ 0 & a_{22} & 0 \\ 0 & a_{32} & 0 \\ a_{41} & 0 & 0 \\ 0 & 0 & 0 \end{bmatrix}$$

The orthogonal complement $\alpha_\perp$, multiplied by the cumulated residuals, gives the following two common trends:

$$\begin{bmatrix} \sum_{i=1}^{t} \varepsilon_{i3} - \frac{a_{32}}{a_{22}} \sum_{i=1}^{t} \varepsilon_{i2} \\ \sum_{i=1}^{t} \varepsilon_{i5} \end{bmatrix}$$

where the second is still given by the cumulated errors of final

Table 5: Constancy test

	mean	variance	joint
ECM_1	0.76*	0.72189*	1.19442*
ECM_2	0.47	1.78549**	2.48921**
ECM_3	0.59*	1.93799**	2.76475**

domestic demand but the first trend is a linear combination of shocks to the bond yield and the return on money only.

It is important to notice that the coefficient of the first equilibrium relationship in the equation of the inflation rate is definitely different from zero. This contrasts with some recent results (see Rinaldi and Tedeschi, 1996) where evidence is found of long run exogeneity of all the variables with respect to the weight and the coefficients of money demand in the equation of real money supply[15].

8.4 Constancy test on the long-run equilibria

To assess the constancy of each cointegration vector, we have performed a univariate analysis. Looking at the graph of the three identified long-run equilibria (fig.3) we can see, in fact, that their variability clearly changes from the 70's to 80'. Hansen's stability test (Hansen, 1992) highlights this pointing out that the instability lies mainly in the variance of the second and the third relationship and less in the mean of these equilibria (Table 5).

Some graphical instability tests are also reported on fig.4,5,6[16]. As

[15]Rinaldi and Tedeschi test that the first relationship, which corresponds to money demand, is not present in the dynamics of the other variables. They find some evidence (not too strong, though) that this could be the case. The same test performed in our system rejects strongly this hypothesis. We must notice, however, that results are not strictly comparable as the time span is different.

[16]The recursive graphs are calculated for $t = M...T$. The first graph shows the 1-step residuals $y_t - X_t'\widehat{\beta}_t$ inside the bands $0 \pm 2\widehat{\sigma}_t^2$. If the residuals lie outside the bands then either there is a presence of an outlier or of parameter non-constancy. The third graph shows 1-step F-tests (1-step Chow tests). The following statistic is calculated:

we can see from the first three graphs of each figure, the three relationships show signs of instability around the end of 1974. This phenomenon does not seem to be caused by a permanent shift in the relationships but rather to the presence of a temporary movements in the coefficients around that date due to the consequences of the first oil shock[17].

To investigate further this phenomenon, we split the sample into two subsamples, one covering the period 1970-1982, the other the period 1983-1994. With reference to this split we perform a variance-covariance LR test of the hypothesis of a constant variance-covariance matrix in the two periods against the alternative of a change. We use Box's version of the LR test (Box, 1949). Under the null the statistic is distributed as a $\chi^2(\frac{p(p+1)}{2})$ where p is the number of variables in the system:

Box's $LR = 78.25207$

$\chi^2_{0.05}(15) = 7.26$

$\chi^2_{0.01}(15) = 5.23$

The hypothesis of constancy is clearly rejected. A possible interpretation of the difference in variance is represented by a shift in the coefficients $\alpha's$ from 70's to 80's. However, a preliminary split sample cointegration analysis has not led us to precisely identify this phenomenon in the data and further research will, thus, be needed to better model such a break in variability.

$\frac{(RSS_t - RSS_{t-1})(t-k-1)}{RSS_{t-1}} \overset{H_0}{\sim} F(1, t-k-1)$

The fourth graph shows the break point F-tests (N↓-step Chow tests). It is based on the following statistic:

$\frac{(RSS_T - RSS_{t-1})(t-k-1)}{RSS_{t-1}(T-t+1)} \overset{H_0}{\sim} F(T-t+1, t-k-1)$

Finally, the fifth graph reports the forecast F-tests (N↑-step Chow tests). The statistic is the following:

$\frac{(RSS_t - RSS_{M-1})(t-k-1)}{RSS_{M-1}(t-M+1)} \overset{H_0}{\sim} F(t-M+1, M-k-1)$

(see Doornik and Hendry, 1994)

[17]The instability around this date is probably associated with a sharp rise in inflation due to the first oil shock which led to a rise in interest rates and caused important capital losses to savers (see Caranza and Cottarelli, 1987)

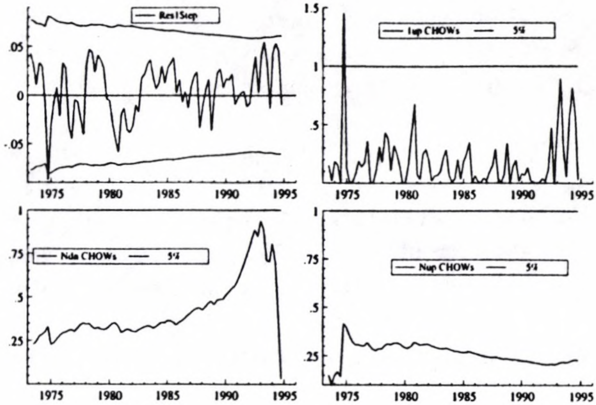

Figure 4: Constancy test on the first cointegration vector

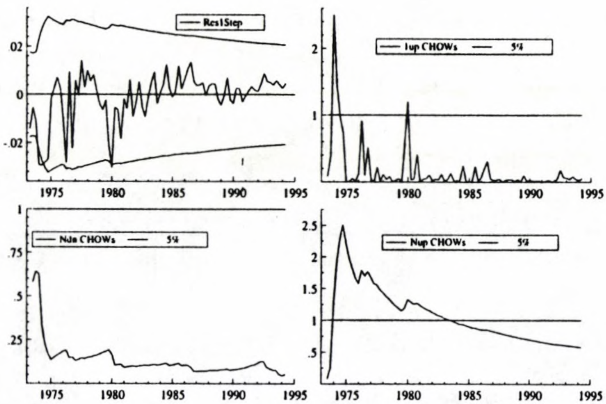

Figure 5: Constancy test on the second cointegration vector

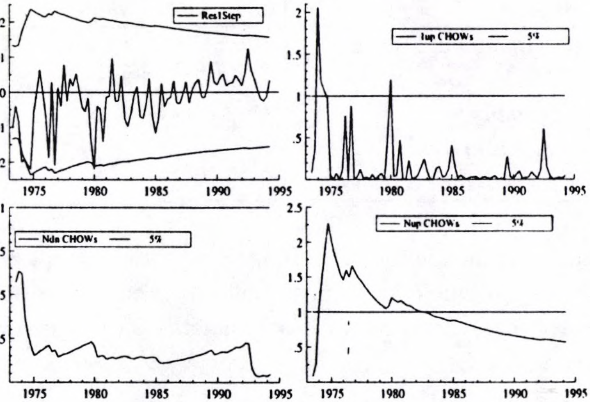

Figure 6: Constancy test on the third cointegration vector

9 From the system to the model

The estimated VAR system is reparameterized in equilibrium correction form (see (2)) and, through successive steps, reduced to a parsimonious representation (PVAR).

The sequential procedure, that starts from the general system to reach a reduced but still congruent configuration, has many advantages (see Mizon, 1995). Some of the methodological ones include the possibility of having a specific direction of research (that avoids thinking about all the directions in which the simple model could be expanded) as well as that of avoiding to adopt the alternative hypothesis when the null has been rejected (see Hendry, 1995). Furthermore, the test for overidentifying restrictions allows us to judge whether the model encompasses the general system from which it is derived and can then be considered a valid representation of the data generating process of the modeled series (see Hendry and Mizon, 1993; Mizon, 1984). This is important, from an economic point of view because it can then be used as a testing ground

Table 6: Diagnostics of the parsimonious VAR

Diagnostics	statistic	value	*pvalue*
Vector portmanteau 10 lags	126.8		
Vector $AR(1-5)$	$F(80,215)$	$= 0.63172$	[0.9908]
Vector Normality	$\chi^2(8)$	$= 5.4765$	[0.7056]
Likelihood Ratio Test for over identifying restrictions	$\chi^2(56)$	$= 53.3259$	[0.5767]

for alternative economic theories. In fact, a particular interpretation of the underlying economic mechanism that generates the variables which imposes a set of restrictions on the parameters can be tested against the PVAR.

Turning to our estimation, having observed in the previous section long run weak exogeneity of real final domestic demand, we decided to open the system and condition on this variable. The resulting parsimonious VAR, which is not fully reported here, is a congruent representation as it is shown in Table 6.

On the basis of the PVAR, a simultaneous equation model (SEM) is formulated and tested. In Tables 9 and 10 we report the estimates for the first four variables[18]. Together with zero restrictions on the parameters we have also tested for the equality in absolute value of the coefficients of the $\Delta^2 p_{t-3}$ and $\Delta^2 p_{t-4}$ in the equation of the own rate on M2. The test result shows acceptance of this hypothesis:

$$\textit{Wald test} \quad \chi^2(1) = 0.023332 \qquad \textit{pvalue} = 0.8786$$

Notice that, exploiting the validity of the restriction, we have reparameterized the model substituting $\Delta^2 p_{t-3}$ and $\Delta^2 p_{t-4}$ with $\Delta^3 p_{t-3}$.

This formulation allows to point out that the return on M2 depends actually on the second difference of the inflation rate, that is on the acceleration rate of inflation lagged three times.

[18]The variables d1, d2, and d3 are the centred seasonal dummies described in par.8.1.

The system diagnostic statistics do not indicate any misspecification in the model (see Table 7). However, the single equation statistics, which are omitted here, show some very small autocorrelation (significant at 5%) for the third and the fourth equation. Nevertheless, we still consider it as a valid representation of the underlying process that describes the series and a good balance between the need of having a congruent model and that of parsimony in the number of parameters.

An important feature to notice is the significance of the equilibria in the individual equations. Real money supply reacts to both disequilibria in money demand and in the relationship between the opportunity cost and inflation, while the own rate of return on money reacts only to Fisher's parity.

It is important to notice that the bond yield does not react to any disequilibria and it is, thus, long run weakly exogenous. In terms of stochastic trends (see also par.8.3) this means that the system is driven by shocks to real expenditures and also by shocks to the bond rate:

$$\begin{bmatrix} \sum_{i=1}^{t} \varepsilon_{i3} \\ \sum_{i=1}^{t} \varepsilon_{i5} \end{bmatrix}$$

The evidence found here is consistent with a bond yield determined outside the system, by the foreign sector and by the dynamics of the public sector deficit. During the 80's this, in fact, drives up the level of the yield on government bonds and causes a continous appreciation of the exchange rate which is only partly counterbalanced by devaluations of the currency and leads to 1992 exchange rate crisis. The important role of the yields is discussed also in Juselius (1997).

Finally, looking at the impact of excess money supply on inflation, what emerges from this estimation is a strong effect of the monetary conditions on the rate of change in prices. This result contrasts, for instance, with what is found by Rinaldi and Tedeschi (1996) where money demand does not have a significant weight on the inflation equation, and by Bagliano (1996) who initially excludes inflation from the joint

Table 7: Diagnostics of the model

Diagnostics	statistic	value	*pvalue*
Vector portmanteau 10 lags	129.93		
Vector $AR(1-5)$	$F(80, 219)$	$= 0.6217$	[0.9927]
Vector Normality	$\chi^2(8)$	$= 3.2651$	[0.9166]
Likelihood Ratio Test for over identifying restrictions	$\chi^2(57)$	$= 48.5939$	[0.7783]

modeling and from the long run equilibria on the basis of its stationarity and then finds support for strong exogeneity of this variable. The sample size is, however, shorter than the one considered here and it suggests once more the possibility of a change in the exogeneity properties of certain variables when passing from the 70's to the 80's. A split sample analysis performed on the same period (see Juselius and Gennari, 1998), however, seem to indicate that what changes between the two periods is mainly the set of long run equilibria.

9.1 Stability analysis and forecasting properties

Recursive estimation of the system allows the detection of possible parameter non-constancy. In the graphical tests that we report on fig.7 the first three pictures show 1-step residuals with $\pm$ twice their standard errors, while the last three are the 1-step, N↓-step and N↑-step Chow tests already described in the previous section. As we can see, no signs of instability emerge from this battery of tests.

Finally, we have tested the forecasting properties of the model (see also fig.8) using the last four observations of the dataset, i.e. from the first to the last quarter of 1995. The results are reported in Table 8[19]:

[19]The values in square brackets are prob.values. $\widehat{\Omega}$ is the simple variance-covariance matrix of the residuals while V[e] is a variance-covariance matrix that takes into account both innovation and parameter uncertainty (see Doornik and Hendry, 1997, p.198).

Table 8: Parameter constancy forecast test

1-step (ex post) forecast analysis 1995 (1) to 1995 (4)						
with $\widehat{\Omega}$	$\chi^2(16)$	24.085	[0.0877]	F(16,77)	1.5053	[0.1196]
with V[e]	$\chi^2(16)$	21.262	[0.1686]	F(16,77)	1.3289	[0.2020]

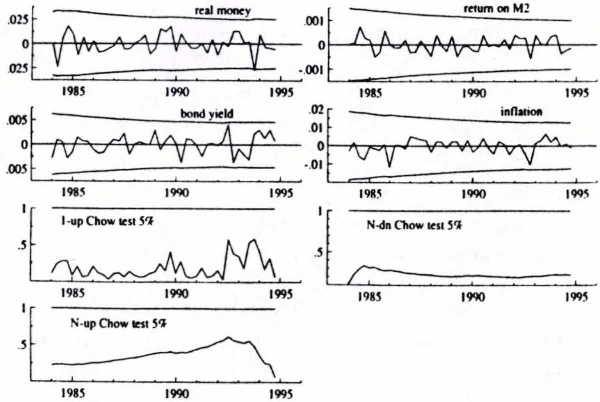

Figure 7: Recursive analysis on the final model

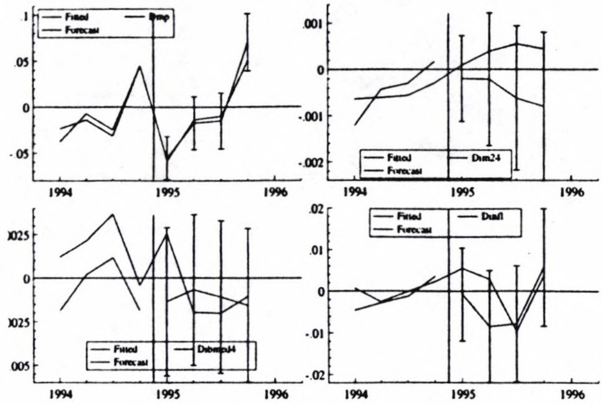

Figure 8: Dynamic Forecasts with final model

10 Conclusions

In this paper we have estimated a money demand relationship for Italy starting from the beginning of 70's until 1994. The period is characterized by relevant changes in the money market, and by an important process of financial innovation mainly driven by an increasing government debt which needed to be efficiently managed.

The empirical analysis has pointed out the importance of accounting for this process in modeling the monetary sector in order to better identify the long-run equilibria amongst the relevant variables.

The learning about new financial instruments has been approximated by a logistic smooth transition function estimated in a preliminary stage with a univariate model, and then introduced in a vector autoregression for cointegration analysis using Johansen's procedure. This analysis has allowed us to identify one of the cointegration vectors as a money demand relationship which is a function of the opportunity cost and the scale variable. The other identified long-run equilibria are relationships amongst the yields and inflation rate and can be connected to term structure of interest rates and to the Fisher's parity.

With a process of simplification, the estimated system has then been reduced to a model which has been found to encompass the general system. Results obtained from this last model show, in particular, the importance of cumulated shocks to the bond yield and to real expenditures in driving the dynamics of the system and of excess money supply in determining the dynamics of the inflation rate. The result is in contrast with what has been found, for instance, by Bagliano (1996) and Rinaldi and Tedeschi (1996). However, their previous work used a sample which ranged from the 80's to the beginning of 90's, and thus suggests the possibility of a change in the exogeneity properties of certain variables from the 70's to the 80's.

This is confirmed also by a stability analysis of the three vectors which has highlighted the presence of changing variability between the 70's and the 80's, probably due to a different adjustment of the variables to the long-run equilibria in the two subperiods. This calls for further research in the direction of modeling changing behaviour of the adjustment parameters, particularly of a nonlinear type, and on the investigation over the consequences of the omission of such a phenomenon.

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Table 9: First two equations of final model

$\Delta(m-p)_t$			
variable	coeff	std.er.	t-prob
$\Delta(m-p)_{t-3}$	0.2267	0.0690	0.0015
$\Delta(m-p)_{t-4}$	0.3595	0.0704	0.0000
$\Delta^2 im_{t-3}$	7.0209	2.0084	0.0008
$\Delta^2 im_{t-4}$	-3.7455	1.7961	0.0403
Δib_{t-1}	-2.9432	0.5614	0.0000
Δib_{t-2}	-1.1130	0.5489	0.0460
Δib_{t-3}	-1.1621	0.6405	0.0735
$\Delta^2 p_{t-2}$	-0.4446	0.1700	0.0107
$\Delta^2 p_{t-3}$	-0.5473	0.1757	0.0026
$\Delta(x-p)_t$	0.7486	0.1264	0.0000
$\Delta(x-p)_{t-1}$	-0.3124	0.1132	0.0072
$\Delta(x-p)_{t-3}$	-0.2047	0.1188	0.0886
$ecm1_{t-1}$	-0.3388	0.0516	0.0000
$ecm3_{t-1}$	0.8439	0.1949	0.0000
i1974p4	-0.0296	0.0138	0.0349
i1986p2	0.0058	0.0085	0.4978
d1	0.0139	0.0051	0.0079
d2	0.0020	0.0038	0.6035
d3	0.0054	0.0048	0.2587
constant	0.14194	0.0200	0.0000

Δim_t			
variable	coeff	std.er.	t-prob
$\Delta(m-p)_{t-1}$	0.0123	0.0027	0.0000
$\Delta(m-p)_{t-2}$	0.0070	0.0024	0.0059
Δim_{t-1}	0.6854	0.0888	0.0000
Δim_{t-1}	-0.4725	0.0935	0.0000
Δib_{t-1}	0.1504	0.0297	0.0000
Δib_{t-3}	0.0734	0.0292	0.0140
$\Delta\Delta^2 p_{t-3}$	-0.0113	0.0037	0.0033
$\Delta(m-p)_t$	0.0116	0.0048	0.0196
$\Delta(x-p)_t$	-0.0228	0.0072	0.0022
$ecm2_{t-1}$	-0.0547	0.0083	0.0000
i1974p4	0.00006	0.0005	0.9094
i1986p2	-0.0012	0.0003	0.0008
d1	-0.0008	0.0002	0.0002
d2	-0.0004	0.0001	0.0134
d3	-0.0002	0.0001	0.0413
constant	-0.0009	0.0001	0.0000

Table 10: Last two equations of final model

Δib_t				$\Delta^2 p_t$			
variable	coeff	std.er.	t-prob	variable	coeff	std.er.	t-prob
$\Delta(m-p)_{t-4}$	0.0125	0.0098	0.2051	$\Delta(m-p)_{t-4}$	0.1001	0.0341	0.0044
Δib_{t-4}	-0.1656	0.0914	0.0737	Δim_{t-1}	4.1147	1.041	0.0004
$\Delta^2 p_{t-1}$	-0.0601	0.0317	0.0619	Δib_{t-2}	-0.7265	0.3246	0.0281
$\Delta^2 p_{t-2}$	-0.0915	0.0320	0.0055	$\Delta^2 p_{t-1}$	-0.6134	0.0849	0.0000
$\Delta(x-p)_{t-1}$	0.0250	0.0199	0.2114	$\Delta^2 p_{t-2}$	-0.6190	0.1105	0.0000
$\Delta(x-p)_{t-3}$	0.0345	0.0210	0.1051	$\Delta^2 p_{t-3}$	-0.5019	0.1100	0.0000
Δim_t	1.5685	0.3201	0.0000	$\Delta^2 p_{t-4}$	-0.1671	0.0692	0.0182
i1974p4	-0.0008	0.0020	0.7020	$\Delta(x-p)_t$	-0.3519	0.0691	0.0000
i1986p2	-0.0021	0.0013	0.1237	$\Delta(x-p)_{t-1}$	0.3907	0.0584	0.0000
d1	-0.0005	0.0005	0.3489	$\Delta(x-p)_{t-3}$	0.1127	0.0605	0.0662
d2	0.0006	0.0005	0.2663	$\Delta(x-p)_{t-4}$	0.2892	0.0707	0.0001
d3	-0.0007	0.0008	0.3769	ecm1_{t-1}	0.1437	0.0263	0.0000
constant	-0.0005	0.0003	0.1176	i1974p4	-0.0021	0.0065	0.7370
				i1986p2	-0.0050	0.0042	0.2372
				d1	-0.0095	0.0028	0.0013
				d2	0.0032	0.0022	0.1436
				d3	-0.0051	0.0027	0.0646
				constant	-0.0590	0.0100	0.0000

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