Economic Notes by Banca Monte dei Paschi di Siena SpA, vol. 25, no. 3-1996, pp. 425-464

Money Demand in a Multivariate Framework: A System Analysis of Italian Money Demand in the '80s and Early '90s

FABIO C. BAGLIANO *

Money demand is studied in a multivariate framework, so as to explicitly address the problem of multiple long-run relations, which is overlooked in single-equation estimates. The adopted methodology, combining cointegration analysis with traditional structural modelling, has two distinctive features: (i) it makes use of formal testing of long-run structural economic hypotheses in the context of a cointegrating VAR; (ii) the dynamic, short-run, adjustment of the system is specified in a way consistent with the proposed interpretation of the long-run equilibrium. This empirical strategy is illustrated with an application to Italian data for the eighties and early nineties.

(J.E.L.: C32, E41).

Introduction

Much of the recent empirical work on money demand has focused on several serious episodes of structural instability of the estimated equations, especially for the US. Various authors have put forward *non-structural* interpretations of the conventionally estimated money demand functions obtained through single-equation empirical models. In particular, two explanations of this kind have received attention. On the one hand, interpreting money demand equations as inverted price equations (Gordon [1984], Carr, Darby and Thornton [1985]), instability episodes may be due

* Dipartimento di Scienze Economiche e Finanziarie "G. Prato", Università di Torino, Corso Unione Sovietica 218bis, 10134 Torino (Italy); e-mail: bagliano@econ.unito.it

I would like to thank Charlie Bean, Carlo Favero, Roberto Rinaldi, Roberto Tedeschi, Simon Wren-Lewis, two referees and seminar participants at Bocconi University for useful comments and suggestions. Paolo Angelini and Roberto Rinaldi (Bank of Italy) kindly provided part of the data set used in the empirical analysis. Financial support from MURST (40% and 60%) is gratefully acknowledged.

to the omission of supply-side variables (e.g. oil prices) directly affecting the price level (however, Fischer and Nicoletti [1993] have effectively crticized this argument on the basis of weak exogeneity tests in a cointegration framework). On the other hand, due to forward-looking behaviour on the part of money holders, the estimated money demand parameters may well be complicated convolutions of structural elasticities, describing agents' behaviour, and expectations parameters, reflecting the information set available to agents and the particular way in which expectations are formulated. Such estimated parameters may then display instability over time only because the process generating expectations has altered, with no change in the underlying structural money demand elasticities (Cuthbertson and Taylor [1990]). Following a different line of research, Baba, Hendry and Starr (1992) attribute the detected instability to the omission of relevant variables and to dynamic misspecification. Once some measure of risk and return on long-term bonds is included in an otherwise conventional money demand function and the equation's dynamic structure is more carefully determined, instability during the well-known "missing money" (1974-76) and "great velocity decline" (1982-83) episodes in the US is eliminated.

Applying the recently developed cointegration techniques (Johansen [1988], Johansen and Juselius [1990]) to US data, several studies have provided support for the existence of a *long-run* money demand function with stable income and interest rate elasticities over extended time periods and at different data frequencies (Hafer and Jansen [1991], Hoffman and Rasche [1991], Stock and Watson [1993], Hoffman, Rasche and Tieslau [1995]; Muscatelli and Spinelli [1996] analyse Italian historical data on money demand). However, when money demand is viewed in the broader context of a *system* of variables, including for example income, interest rates and inflation, the possibility arises of the existence of *multiple long-run relations* among the variables. If this is the case, conventional single equation analyses provide estimates of the long-run money demand that are instead combinations of the multiple relations linking the series under study. The cointegration techniques proposed by Johansen, Juselius and others, yielding tests for the number of long-run relations in a system of variables and estimates of the form of such relations, have been extensively used to face this problem (a recent application to US data is the joint analysis of money demand and the interest rate term structure by Rasche [1994]).

In line with the above view, our aim here is to specify a structural multivariate model of the long-run and short-run interrelationships among the variables usually involved in the analysis of money demand. The adopted approach has two distinctive features: (*i*) it makes use of formal testing of long-run structural economic hypotheses by means of the likelihood ratio tests developed by Johansen and Juselius (1992, 1994) in the context of a cointegrated *VAR* system; (*ii*) subsequently, a simultaneous structural model is specified, with a short-run dynamics consistent with the economic interpretation of the long-run equilibrium path of the system. This model is then tested against the (reduced-form) cointegrated *VAR*.

We apply our approach to the analysis of the recent (pre-ERM crisis) behaviour of Italian money demand, since various studies have highlighted the presence of multiple long-run relations among money balances, income and interest rates (Muscatelli [1991], Bagliano and Favero [1992]), without formally testing structural hypotheses on their economic nature. The focus on the estimates of the long-run features of the data requires stability of the underlying economic relations; the intense process of financial innovation occurred in Italy particularly in the late seventies, potentially causing changes of the long-run relations between money demanded, income and interest rates, motivates our choice of a sample period starting only in the early eighties.

The next section further motivates our approach, discussing its relation with the existing literature. The empirical methodology adopted and the results obtained are reported in section 2. The final section summarizes the main conclusions.

1. *Motivation and Related Literature*

Our analysis illustrates the need for the adoption of a multivariate approach to the empirical study of money demand. The fact that money demand functions are part of a larger system of equations describing the complex interrelationships among money balances, income, interest rates and inflation is often recognized also in the context of single-equation modelling, motivating the use of instrumental variables techniques. However, besides the need for correcting for simultaneity bias, there are other reasons to justify an explicit multivariate approach, even though the interest is in modelling only one economic function. In fact, even when simultaneity is not a relevant problem, overlooking the presence of more than one long-run, cointegrating relation may lead to serious misinterpretations of the long-run properties of agents' behavior and also to mis-specifications of the short-run dynamic adjustment towards equilibrium[1].

[1] Banerjee, Dolado, Galbraith and Hendry (1993), Campbell and Perron (1991) and Ericsson (1992) provide detailed accounts of cointegration theory with extensive bibliographies.

A simple example can illustrate this point. Let us consider four variables (lower case letters denote logarithms): real money balances ($m-p$), real expenditure (y), the own yield on money (R^m) and an alternative interest rate (R^b). We assume that the following two long-run relations hold:

$$m - p = \alpha y \tag{1}$$

$$R^m = \gamma R^b \tag{2}$$

The first equation implies that money demand is determined by expenditure only, with no long-run interest rate effects, whereas the second posits a long-run relation between the two rates, possibly determined by the banking sector's behaviour in setting the deposit rate. Now, let the short-run dynamics of the system be determined according to the following four equations:

$$\Delta(m-p)_t = a_1\,\Delta(m-p)_{t-1} - a_2\left[(m-p) - \alpha y\right]_{t-1} + a_3\,(R^m - \gamma R^b)_{t-1} + u_{1t} \tag{3}$$

$$\Delta y_t = b_1\,\Delta y_{t-1} + b_2\left[(m-p) - \alpha y\right]_{t-1} + u_{2t} \tag{4}$$

$$\Delta R^m_t = c_1\,\Delta R^m_{t-1} - c_2\,(R^m - \gamma R^b)_{t-1} + u_{3t} \tag{5}$$

$$\Delta R^b_t = d_1\,\Delta R^b_{t-1} + u_{4t} \tag{6}$$

Both money balances and expenditure react to past deviations of money demand from the equilibrium (long-run) relation (1). Also the interest rate on money displays error-correcting behaviour, since the relevant disequilibrium term enters equation (5). Moreover, the same interest rate error-correction term enters the money balances equation, indicating that although in the long-run money demand is independent of interest rates, deviations of interest rates from their equilibrium path may affect the short-run dynamics of money balances. In order to focus on the problems caused by the presence of multiple long-run relations, no simultaneous term is included. The additional assumption of independent disturbance terms in (3)-(6) allows the estimate of a single money demand equation not to suffer from simultaneous equation bias. If a single-equation money demand analysis is performed on the data, a likely outcome, observationally equivalent to (3), is the following:

$$\Delta(m-p)_t = \delta_1\,\Delta(m-p)_{t-1} - \delta_2\,(m-p)_{t-1} + \delta_3\,y_{t-1} + \delta_4\,R^m_{t-1} - \delta_5\,R^b_{t-1} + \varepsilon_t \tag{7}$$

The estimated long-run solution, obtained from the terms in levels in (7), may be erroneously interpreted as a money demand function with non-zero interest rate elasticities (δ_4/δ_2 and δ_5/δ_2 respectively for R^m and R^b). A system analysis is necessary in order to detect the existence of two distinct long-run relations, since the presence of the disequilibrium terms $(m-p)-\alpha y$ and $R^m - \gamma R^b$ in (4) and (5) imposes (testable) cross-equation restrictions on the system parameters. These restrictions, either implied by some economic theory or suggested by unrestricted estimation of (3)-(6), with the terms in levels capturing the long-run features of the data, may then be imposed and tested on the whole system, providing information that the one-equation money demand analysis is bound to overlook.

One strand of the recent applied econometric literature, following a tradition dating back to the Cowles Commission, has focused on the formulation and estimation of linear dynamic simultaneous systems starting from a general reduced form (a vector autoregression [*VAR*]). Empirical observation and *a priori* economic theory may then be used to obtain identification of a *structural* simultaneous equations model. The emphasis is placed on the formulation (using mis-specification and parameter stability tests) of a data-coherent reduced form system, providing a valid framework for evaluating structural economic hypotheses by means of encompassing tests (Hendry, Neale and Srba [1988], Monfort and Rabemananjara [1990], Clements and Mizon [1991] and Hendry and Doornik [1994]). The non-stationary nature of most macroeconomic time series requires the adoption of appropriate methodologies for system estimation and inference. Johansen (1988, 1991) and Johansen and Juselius (1990) addressed the problem of estimating the long-run equilibrium relations (cointegrating vectors) among non-stationary variables in a multivariate context, devising a procedure to test for the number and form of such relations. The information so obtained on the long-run properties of the data may then be incorporated in the system specification strategy, in order to reach a complete characterization of the short-run dynamics of the variables, adjusting towards their equilibrium path. This kind of cointegrated *VAR* is then viewed as the appropriate specification of the system's reduced form, capturing the long-run features of the series, from which to start the process of formulation and testing of alternative structural (simultaneous) models (Clements and Mizon [1991], Hendry and Mizon [1993], Chow [1993], Hendry and Doornik [1994])[2].

[2] Other authors prefer to apply traditional *VAR* techniques (impulse response functions and forecast error variance decompositions) to the cointegrated reduced form system, avoiding the imposition of (over)identifying structural restrictions onto the data (King, Plosser, Stock and Watson [1991]).

The present paper follows this approach, combining Johansen's long-run analysis with the structural modelling strategy proposed by Hendry, Mizon and Chow. The adopted procedure is divided into two subsequent steps. First, we study the *long-run* behavior of the data (money balances, income, interest rates and inflation), estimating the number and form of the cointegrating vectors; at this stage, specific hypotheses are formally tested in order to provide an economically meaningful interpretation for the detected long-run equilibrium relations. Second, a dynamic simultaneous system is specified, including the disequilibrium (error-correction) terms constructed from the estimated cointegrating vectors and embedding the long-run structural economic hypotheses tested in the preceding step. At this stage, structural hypotheses on the *short-run* dynamics of the system are formulated. In addition to the contemporaneous relations suggested by economic theory, also some hypotheses on the responses of the endogenous variables to deviations from the equilibrium path may be specified. The previously tested structural assumptions on the long-run behavior of the system may suggest a pattern of error-correcting responses of the variables consistent with the economic interpretation of the system's equilibrium path. For example, the economic nature of the series may suggest that some variables should display a stronger tendency to react to disequilibrium than others. Furthermore, the short-run dynamics of some variables may be influenced by more than one error-correction term associated with the long-run equilibrium relations of the system. The resulting restrictions may finally be tested against the system's reduced form.

Though in principle the procedure outlined here may not be capable of settling conclusively the observational equivalence problem illustrated at the beginning of this section, the system approach has two clear advantages over single-equation modelling: (*i*) the issue of multiple long-run equilibrium relations in the system is directly addressed, and (*ii*) the estimated short-run dynamic adjustment of the endogenous variables is consistent with the economic nature of the system's equilibrium path.

A similar estimation methodology is applied by Clements and Mizon (1991) to the study of wage and price determination in the U.K. over the period 1965-1989. Multiple cointegrating vectors are found by Hendry and Mizon (1993) and Hendry and Doornik (1994) in estimation of a small monetary model for the U.K. (1963-1989). Here two valid long-run relations are found: a demand for money function and an excess aggregate demand equation (linking the deviation of output from trend to inflation and the interest rate). In the structural simultaneous model real money balances react in an error-correcting way to deviations from the long-run money demand whereas excess demand triggers equilibrating responses of expenditure, inflation and – in the Hendry-Doornik version only – the interest rate.

Within a three-variable system, Chow (1993) constructs a simple structural multiplier-accelerator model, with the consumption-income and investment-income ratios capturing the long-run equilibrium of the economy. Here it is explicitly noted that in a system context with m structural equations and $r < m$ cointegrating vectors one "*cannot associate each structural equation with an error-correction mechanism attributable to that equation alone. All of the r* [cointegrating vectors]... *may affect the dependent variable of the ith structural equation... Sometimes a structural equation may have an error-correction term attributable only to an equilibrium relationship among its own variables. Sometimes* [it] *may have no* ...[or] *several cointegrating vectors associated with it*", (p. 110). This point is noted also by Konishi, Ramey and Granger (1993) in analyzing the interrelationships between real and financial variables in the U.S. over the 1960-1991 period. Finally, Johansen and Juselius (1994) have recently applied a similar procedure to macroeconomic data for Australia.

Our empirical investigation follows the spirit of Chow (1993) and Johansen and Juselius (1994), applying a sequential identification process of the long- and short-run structures to monetary data for Italy. Previous efforts in modelling Italian money demand behaviour are due to Bagliano, Favero and Muscatelli (1991) and Muscatelli (1991). In particular, Muscatelli (1991), using quarterly data for the period 1966-1984, explicitly recognized the need for a multivariate approach in the presence of multiple cointegrating vectors. Two long-run relationships between money balances, income, and interest rates were estimated, both apparently interpretable as money demand functions, though with widely different elasticities. No structural hypotheses were tested on these vectors, which were included in the structural system as originally estimated. One of the error-correction terms was found to enter the equations for money balances and inflation, the other causing adjustments of the money yield and of the interest rates on alternative assets. This pattern of short-run responses of the endogenous variables to disequilibrium was not given a structural economic interpretation. The methodology we adopt in this paper differs from that in Muscatelli (1991) in at least two respects[3]: (*i*) we formulate and test explicit structural hypotheses on the nature of the cointegrating vectors detected, and (*ii*) we specify a pattern of adjustment of the endogenous variables consistent with the economic interpretation put forward for the long-run equilibrium, testing the resulting restrictions on the dynamics of the system.

[3] An analysis of money demand behaviour closely related to our approach has recently been provided by Rinaldi and Tedeschi (1995).

2. *Empirical Methodology and Results*

2.1. *Setting up the VAR*

The first issue addressed here is the choice of the endogenous variables to be modelled in the system analysis. This amounts to specifying the long-run determinants of money demand in the period under consideration (1983-1991). Our choice is guided by basic money demand theory, suggesting a role for a scale variable, the yield on alternative assets, the own return on the interest-bearing components of the relevant monetary aggregate and, perhaps, the inflation rate. We begin the data analysis by investigating the integration properties of the following variables (lowercase letters denote logarithms): nominal M2 money balances (m), the consumer price index (p), real money balances ($m-p$), total final expenditure (real GDP plus net real imports, y), the after-tax yield on Treasury bills averaged over three-, six-, and twelve-month maturities (R^b), and the after-tax own return on M2, obtained as a weighted average on the various components of the monetary aggregate (R^m). All series are monthly, from 1983(1) to 1991(12), and nominal money and expenditure are seasonally adjusted[4]. The results of a battery of Augmented Dickey-Fuller (*ADF*) tests on these variables show that all variables may be considered $I(1)$, with some evidence of a deterministic trend only for the price level.

At this stage, two modelling choices are made. First, in order to reduce the dimension of the system and aid the economic interpretability of the cointegration results, money balances are included in the *VAR* in real terms, thereby imposing long-run homogeneity of degree one of nominal money balances to the price level (formal support for this assumption will be provided by the cointegration analysis of the next subsection). Second, given the stationary ($I(0)$) behavior displayed over the estimation period, the inflation rate (Δp) is excluded from the *long-run* determinants of money demand. However, a dynamic *short-run* effect on the endogenous variables is allowed by including Δp in the system as an exogenous, conditioning variable. The (weak) exogeneity of the inflation rate for the parameters of interest will be appropriately tested in the following analysis, where also an additional test of the $I(0)$ nature of this series will be performed. The resulting system therefore includes as endogenous variables: real money balances, total final expenditure and the yields on money and on Treasury bills.

Prior to studying the long-run properties of the system, we perform a variable by variable analysis using reduced form models for $\Delta(m-p)$, Δy,ΔR^b and ΔR^m, in order to detect anomalies in their time-series behavior and assess the potential role of additional exogenous variables in each individual equation. To this aim, letting $\mathbf{x}_t = \{(m-p), y, R^b, R^m\}'$, we start from a basic four-lag *VAR* specification, rearranged in order to express the dependent variables in first difference form as follows:

$$\Delta \boldsymbol{x}_t = \boldsymbol{A}_0 \boldsymbol{x}_{t-1} + \sum_{i=1}^{3} \boldsymbol{A}_i \Delta \boldsymbol{x}_{t-i} + \delta \Delta p_t + \boldsymbol{c} + \boldsymbol{u}_t \tag{8}$$

where $\mathbf{A}_0$ and $\mathbf{A}_i$ ($i = 1, 2, 3$) are 4 by 4 matrices, δ is a four-element vector of coefficients, Δp is the inflation rate (included in the basic specification as the only contemporaneous conditioning variable[5]), c is a vector of constant terms and u_t is the vector of residuals. Each equation of the above system is then separately estimated and the residuals tested for normality and serial correlation. The results are reported in the first panel of Table 1. In all equations, huge residual non-normality is detected. For the real money balances and expenditure equations this behavior seems attributable to isolated episodes and two dummy variables (described in the notes to Table 1) are introduced to take care of such outlier observations. As shown in the second panel of the table, the inclusion of these two dummies is sufficient to eliminate residual non-normality from the $\Delta(m-p)$ and Δy equations.

In order to obtain a satisfactory specification for the two interest rate equations, additional dummies (detailed in the notes to Table 1) and also exogenous variables are needed. In particular, to capture the effect of monetary policy actions on market rates, changes in the discount rate ($\Delta DISC$) and in the rate on repurchase operations of the central bank ($\Delta REPR$) are included in the equation for ΔR^b. As shown in the final panel of the table, both policy variables have a statistically significant effect on the Treasury bill rate, much higher for the discount rate. Lagged changes in policy rates have only a small (and statistically not significant) additional effect, suggesting that the transmission of monetary policy

[4] We use the new definition of M2, recently adopted by the Bank of Italy in order to improve the comparability of monetary aggregates with other European Community countries and first employed by Angelini, Hendry and Rinaldi (1994), who also provide the monthly real GDP series. This is obtained by applying the methodology of Chow and Lin (1971) to the quarterly figure using the available monthly industrial production as a "reference series".

[5] The inclusion of Δp in (8), though not relevant to the determination of the system's long-run properties, reduces the number of outliers in the residuals from the money balances and interest rate equations. In all equations, lags of Δp are not statistically significant.

Table 1

Single-equation Analysis

Dependent variable:	$\Delta(m-p)$	Δy	ΔR^b	ΔR^m
Basic specification:				
S. D. of dep. var.	0.550	1.625	0.413	0.122
R^2	0.133	0.421	0.122	0.356
σ	0.512	1.236	0.387	0.098
Norm. χ^2 (2)	15.08 (0.001)	15.20 (0.001)	32.22 (0.000)	56.96 (0.000)
Ser. Corr. F (12)	1.47 (0.15)	1.43 (0.17)	1.90 (0.05)	1.18 (0.31)
With dummies added:				
R^2	0.312	0.510	0.352	0.560
σ	0.456	1.137	0.332	0.081
Norm. χ^2 (2)	4.67 (0.10)	1.05 (0.59)	8.35 (0.01)	4.34 (0.11)
Ser. Corr. F (12)	1.76 (0.07)	1.26 (0.26)	1.37 (0.20	1.08 (0.39)
With dummies and exogenous variables added:				
$\Delta DISC_t$			0.537 (0.094)	
$\Delta REPR_t$			0.101 (0.028)	
$\Delta DISCN_{t-1}$				0.197 (0.028)
$\Delta REPRN_{t-1}$				0.032 (0.011)
R^2			0.597	0.736
σ			0.262	0.063
Norm. χ^2 (2)			1.80 (0.41)	1.38 (0.50)
Ser. Corr. F (12)			1.00 (0.45)	1.17 (0.32)
Funct. Form F	1.79 (0.18)	0.001 (0.99)	1.48 (0.23)	2.81 (0.10)
ARCH (6) F	0.55 (0.77)	0.48 (0.82)	1.52 (0.18)	0.49 (0.81)
Heterosc. F	1.02 (0.31)	3.26 (0.07)	0.04 (0.84)	0.45 (0.50)
Pred. Failure F (12)	2.33 (0.01)	1.87 (0.05)	3.12 (0.001)	1.61 (0.10)

Notes: A. Sample period: 1983(1)-1991(12). Rates of growth are expressed in percentage points, as are interest rates. The *basic specification* is defined by (8) in the text. σ is the standard error of the regression; *Norm.* χ^2 is the Jarque-Bera test for residual normality; *Ser. Corr. F (12)* is the *F*-version of Godfrey's Lagrange Multiplier test for residual serial correlation up to the 12th order; *Funct. Form F* is the *F*-version of the *RESET* test of functional form; *ARCH(6)* is the test for autoregressive conditional heteroscedasticity up to the 6th order in F-form (Engle [1982]); *Heterosc. F* is the *F* test for residual (unconditional) heteroscedasticity (White [1980]) and *Pred. Failure F (12)* is the Chow test for predictive failure over the period 1992(1)-1992(12). Probability values are in parentheses beneath test statistics.

B. The following dummy variables are included in the estimated equations in the central and final part of the table:

i) in the equation for $\Delta(m-p)$, a dummy variable (*DUS*), taking the value of 1 in December 1989 and January 1990 is included in order to eliminate the effect of bank strikes on data reporting (a sharp increase by about 1.6% in money balances in both months; see Angelini *et al.*(1994));

ii) in the equation for Δy a dummy variable (*DU878*) taking the value of 1 only in August 1987 is added to capture a huge 5% drop in expenditure;

iii) in the equation for ΔR^b two dummies are included. The first (*DU877*) is a point dummy in July 1987, when a sharp increase (by about 90 basis points) in interest rates occurred following analogous movements in foreign rates, especially in Japan and the US. The second (*DU8967*), taking the value of 1 in June 1989 and −1 in the following month, reflects a sudden fall of more than 1% in Treasury bills yields in June, completely offset in July, unrelated to developments in foreign financial markets but due to contingencies in Treasury financing needs;

iv) finally, in modelling ΔR^m, two dummies are needed: a point dummy in October 1983 (*DU8310*), when the tax rate on deposits interest was raised to 25% causing a drop of more than 40 basis points in the net return on M2, and a second dummy (*DURM3*) taking the value of 1 in three months (July 1984, September 1985 and January 1988), when large drops of about 20 basis points occurred, the last of which corresponding to a further increase in the tax rate on deposit interest from 25 to 30%.

C. $\Delta DISC$ and $\Delta REPR$ denote changes in the discount rate and in the interest rate on repurchase operations conducted by the Bank of Italy, respectively; $\Delta DISCN$ and $\Delta REPRN$ contain only *negative* changes in the two rates. Standard errors are in parentheses under coefficient estimates.

impulses to key short-term market rates is completed within the month. Monetary policy impulses affect also the own return on money, although the response of R^m is smaller than that of the Treasury bill rate. Moreover, only *negative* changes in the two policy rates ($\Delta DISCN$ and $\Delta REPRN$) are transmitted to money yields and with a one-month lag. Such lagged and asymmetric response of R^m to monetary policy impulses is in accordance with independent evidence on the behaviour of bank deposit rates: e.g. in the Bank of Italy monthly econometric model of the money market (Bank of Italy [1988]) the banks' deposit rate strongly reacts with a one-month lag to negative changes in the discount rate, whereas the response to positive changes is much smaller.

The dummies and the four additional exogenous variables are then included in system (8). It is important to note here that the fairly extensive use of dummy variables to deal with some features of the data (especially interest rates) for which it is difficult to provide a complete explanation, may be justified by the scope of our investigation. In fact, the set of variables analysed is chosen with reference to the main determinants of *money demand* and therefore may well omit various specific determinants of interest rates behavior, responsible for most of the episodes referred to above. The system is now estimated recursively in order to assess its structural stability properties and forecast performance over the January-December 1992 period. All equations show structural stability over the 1987-1991 period (data from 1983 to 1986 are used for initialization) as shown in Figure 1 by means of recursive break-point Chow stability tests. On the contrary, some of them (especially the money balances and the Treasury bill rate equations) display predictive failure over 1992. This finding is confirmed for the whole system by a forecast confidence interval test (a system version of the "predictive failure" Chow test, taking into account both innovation and parameter uncertainty), yielding a value of 2.64 (with a probability value of 0.001). The *EMS* exchange rate crisis of September-November 1992 may have altered the relations among the variables, for example by making the interest rate on alternative assets an imperfect measure of the opportunity cost of holding money. The general

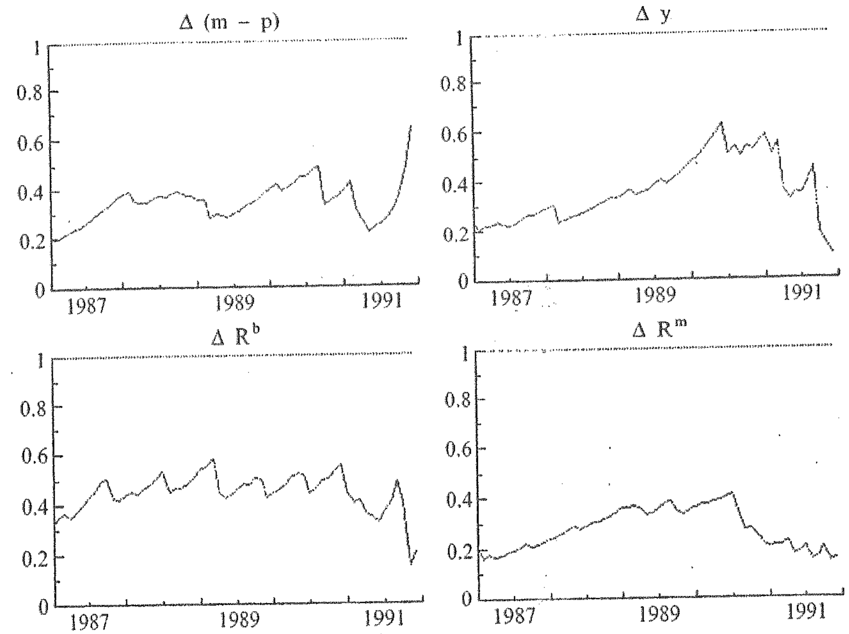


Figure 1: Break-point Chow stability test from recursive system estimation: 1987-1991 (1.0 denotes the 5% crit. value of the test)

uncertainty and the unusual riskiness of alternative financial assets perceived in that period may well be responsible for the underprediction of money balances in the final part of the year. However, also in earlier months some signs of instability are detected, hardly explained by anticipations of an exchange rate crisis, not yet completely reflected in short-term interest rate, leading to a sharp decrease in real balances held by the public. Overall, the system forecast analysis documents the difficulty of extending the estimation period beyond 1991 without introducing additional explanatory variables, possibly augmenting the dimension of the system. Given the main purpose of our investigation, instead of following this route, we chose to end the sample period in 1991 and warn against undue extensions of our results to the more recent period[6].

2.2. *Cointegration Analysis of the Four-variable System*

Having reached an acceptable formulation of the *VAR* system in terms of residual normality and parameter stability, we are now able to apply the maximum likelihood (*ML*) procedure set out by Johansen (1988) and Johansen and Juselius (1990) to test for the presence of multiple cointegrating vectors in a multivariate framework. The procedure yields also an estimate of the valid long-run relationships detected among the variables.

Johansen's methodology is applied to the following system:

$$(9) \qquad \Delta \boldsymbol{x}_t = \boldsymbol{\Pi x}_{t-1} + \sum_{i=1}^{3} \boldsymbol{A}_i \Delta \boldsymbol{x}_{t-i} + \boldsymbol{B}_1 \boldsymbol{d}_t + \boldsymbol{B}_2 \Delta \boldsymbol{w}_t + \boldsymbol{c} + \boldsymbol{\varepsilon}_t$$

where $\mathbf{d}$ and $\Delta\mathbf{w}$ are vectors containing respectively the six dummies and the additional stationary exogenous variables (namely Δp_t, $\Delta DISC_t$, $\Delta REPR_t$, $\Delta DISCN_t-1$ and $\Delta REPRN_{t-1}$), and $\mathbf{B}_1$ and $\mathbf{B}_2$ are conformable matrices. The $\mathbf{\Pi}$ matrix contains all relevant information about the long-run properties of the system. Estimation of (9) is performed under the following assumption of reduced rank of $\mathbf{\Pi}$:

$$(10) \qquad H_1(r): \mathbf{\Pi} = \alpha \beta'$$

where α and β are 4 by r matrices and $r < 4$ is the number (to be estimated) of valid cointegrating vectors in the system. The columns of β

6 Extending the sample period beyond the 1992 ERM crisis to assess the relevance of currency substitution phenomena, as suggested by a referee, is a potentially fruitful direction for further research on money demand (Filosa [1995] provides multi-country evidence on this issue using quarterly data from 1980 to 1992).

form such r vectors, inducing stationarity of the linear combinations of $I(1)$ variables in $\beta' x_{t-1}$, whereas the elements of α are the weights of each cointegrating relation in the equations for the elements of $\Delta \mathbf{x}_t$. Johansen's procedure allows estimation of (9) subject to the reduced rank assumption in (10), yielding estimates of the eigenvalues of the system with corresponding eigenvectors.

The first panel of Table 2 reports the estimated eigenvalues of the four-variable system, together with the values of the Maximum

Table 2

Cointegration Analysis of System: $m-p$, y, R^m, R^b

Eigenvalues:	0.385	0.178	0.109	0.012
Hypothesis:	r = 0	r ≤ 1	r ≤ 2	r ≤ 3
λ_{MAX}	52.5	21.2	12.4	1.4
95% crit. value	27.1	21.0	14.1	3.8
λ_{TRACE}	87.5	35.0	13.8	1.4
95% crit. value	47.2	29.7	15.4	3.8

(r denotes the number of valid cointegrating vectors)

Estimated valid cointegrating vectors (β): $r = 2$

	Original estimates		Normalized on $m-p$		Normalized on R^m	
$m-p$	−0.074	−0.065	−1	−1	0.137	−0.642
y	0.080	0.055	1.085	0.844	−0.149	0.542
R^m	0.540	−0.101	7.296	−1.588	−1	−1
R^b	−0.253	0.050	−3.418	0.761	0.469	0.488

Estimated adjustment matrix (α)

	Original estimates		Normalized on $m-p$		Normalized on R^m	
$m-p$	2.171	−0.394	0.161	−0.026	−1.172	−0.040
y	−2.248	−3.391	−0.167	−0.221	−1.215	−0.344
R^m	−0.197	0.209	−0.015	0.014	0.106	0.021
R^b	−0.483	0.437	0.037	0.028	−0.267	0.044

Estimated long-run matrix (Π) *with reduced rank* $r = 2$

	$m-p$	y	R^m	R^b
$m-p$	−0.135	0.153	1.213	−0.569
y	0.387	−0.367	−0.871	0.401
R^m	0.001	−0.004	−0.127	0.060
R^b	−0.065	0.064	0.222	−0.103

Notes: The estimation period is 1983(1)-1991(12). Cointegration test statistics are obtained by the Johansen (1988) Maximum Likelihood procedure in a four-order *VAR* system including the dummy and exogenous variables listed in notes B and C to Table 2. Critical values for the λ_{MAX} and λ_{TRACE} statistics are tabulated in Osterwald-Lenum (1992).

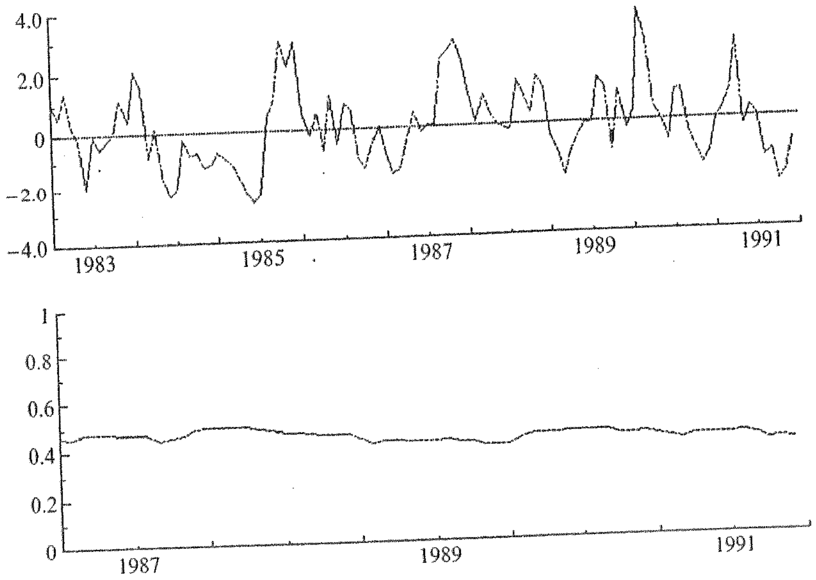


Figure 2: Residuals of the first cointegrating vector adjusted for short-run dynamics and recursive associated eigenvalue

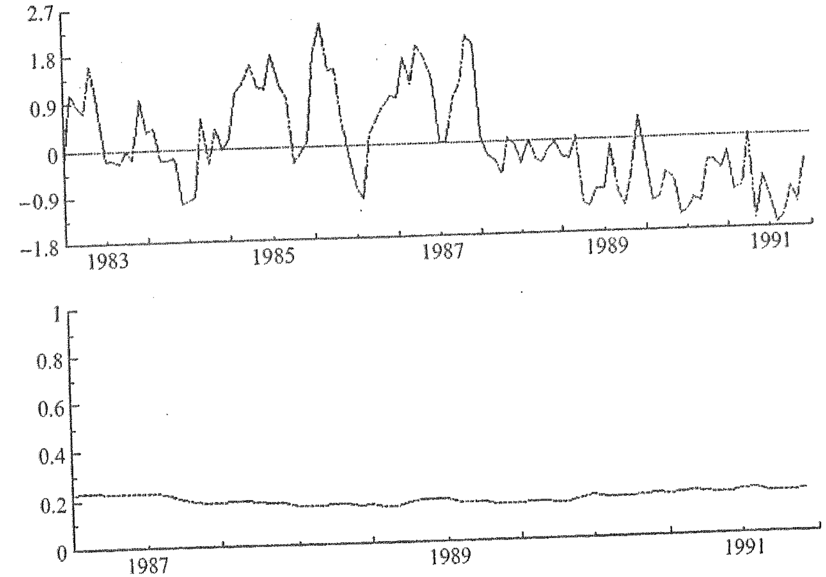


Figure 3: Residuals of the second cointegrating vector adjusted for short-run dynamics and recursive associated eigenvalue

Eigenvalue (λ_{MAX}) and the Trace (λ_{TRACE}) statistics to test for the number (r) of valid cointegrating vectors[7]. Both versions of the test reject the hypothesis of only one cointegrating vector, but do not reject the hypothesis of two such vectors. The elements of the eigenvectors ($\mathbf{v}_1$ and $\mathbf{v}_2$) associated with the two largest eigenvalues of the system are also reported in the table as originally estimated. The linear combinations of the variables in $\mathbf{x}_t$ constructed as $\mathbf{v}_1' \mathbf{x}_t$ and $\mathbf{v}_2' \mathbf{x}_t$ are those most correlated with the stationary part of $\Delta \mathbf{x}_t$ and may be interpreted as the actual deviations of the variables in the system from their long-run equilibrium path. Such deviations are also a function of the short-run dynamics of the system, which may be responsible for their persistence over time. The effect of the short-run dynamics may be eliminated by considering the linear combinations $\mathbf{v}_1' \mathbf{R}_{1t}$ and $\mathbf{v}_2' \mathbf{R}_{1t}$, where $\mathbf{R}_{1t}$ is the vector of residuals from a regression of $\mathbf{x}_{t-1}$ onto $\Delta \mathbf{x}_{t-i}$ ($i = 1, 2, 3$), $\mathbf{d}_t$, $\Delta \mathbf{w}_t$ and a constant. The two resulting cointegrating vectors adjusted for short-run dynamics (shown in Figures 2 and 3 normalized on $m-p$ and R^m respectively) display the required stationary behavior, although the second clearly indicates persistent deviations from the equilibrium path over the final part of the estimation period. Finally, the constancy of the number of valid cointegrating relations throughout the sample is assessed by means of a recursive implementation of the Johansen procedure: the recursive estimates of the two largest eigenvalues obtained (depicted in the bottom part of Figures 2 and 3) show a remarkable stability over the 1987-1991 period[8].

Overall, on the basis of the statistical and graphical evidence presented, we conclude that the four variables in the system are linked by *two* long-run equilibrium relations and proceed under this hypothesis to the estimation of the elements of the α and β matrices. The original estimates provided by the *ML* procedure and reported in Table 2 cannot be given an immediate economic interpretation, since they are obtained from the estimated long-run matrix Π by imposing an arbitrary normalization[9]. Therefore, the estimated columns of β may well be linear combinations (obviously stationary) of the valid cointegrating vectors of

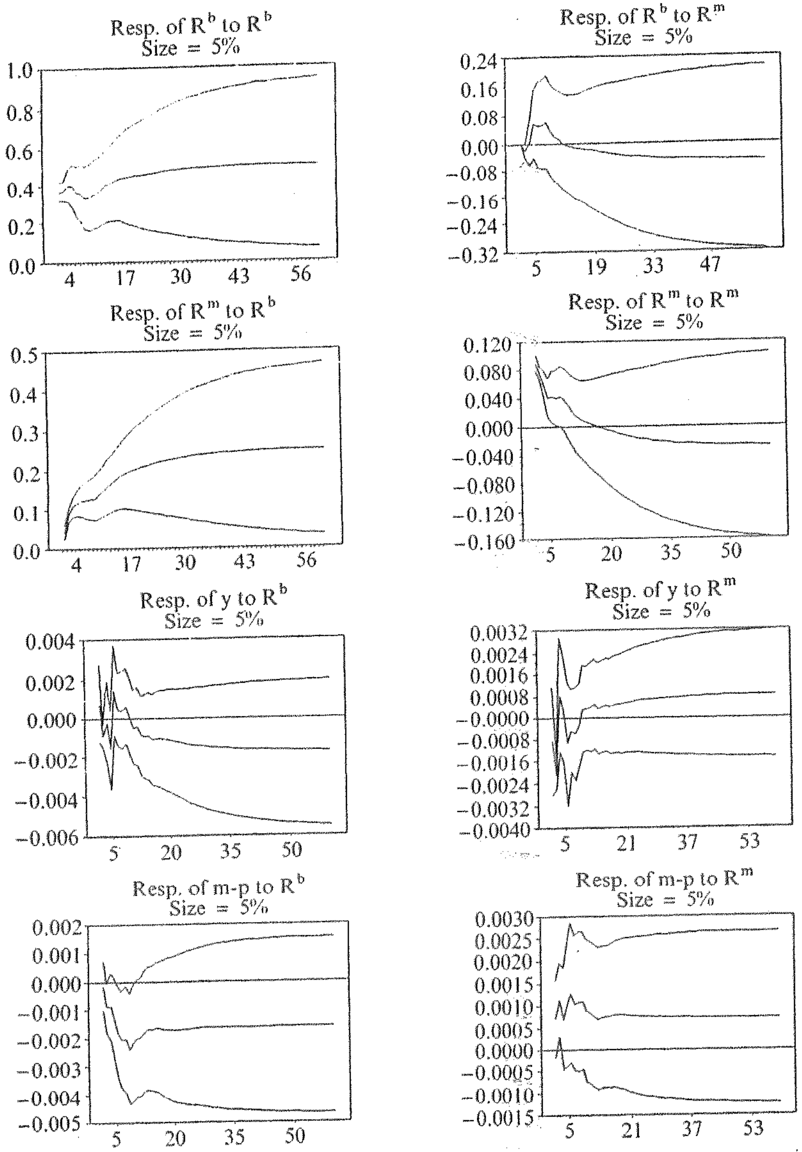
[7] The test based on λ_{MAX} tests the null hypothesis of $r = q - 1$ against the alternative $r = q$, whereas λ_{TRACE} tests $r \leq q - 1$ against $r \geq q$ for all $q = 1, \ldots, n$. Critical values are tabulated in Osterwald-Lenum (1992).

[8] In implementing the recursive procedure the estimated coefficients of the short-run dynamics is kept fixed at the full-sample values (therefore adopting the *R*-representation of the recursion in the terminology of Hansen and Johansen [1992]).

[9] In fact, for any non-singular matrix ξ, we have $\Pi = (\alpha\xi_{-1})(\beta\xi)$. The estimated columns of β may then be rearranged, with corresponding modifications of α, in order to obtain the same matrix Π.

economic interest. In order to aid economic interpretability, we present such estimates under two alternative normalizations. The first is suggested by the main purpose of our investigation, namely the specification and testing of a structural multivariate model of money demand: we then normalize the elements of α and β on real money balances ($m-p$) to assess the possibility of interpreting one of the cointegrating relations as a long-run money demand function. In this respect, the first column of β displays correctly signed coefficients, with an elasticity of real money balances to expenditure close to unity and a negative (positive) long-run response to the alternative (own) return. Viewed from the perspective of a conventional money demand function, the second cointegrating relation displays a plausible value for the expenditure elasticity (0.84) but incorrectly signed coefficients on the two interest rates. The relative magnitudes (in absolute value) of the coefficients on R^m and R^b in the columns of β suggest the second normalization (on R^m) reported in Table 2, confirming this common feature of the two cointegrating relations. Overall, the estimated cointegrating vectors share two common patterns: (*i*) the coefficients on y and $m-p$ are opposite in sign and (after normalization) not very different in magnitude (their ratio ranging, in absolute value, from 0.84 to 1.08); and (*ii*) the coefficients on R^b and R^m are opposite in sign, with an almost identical ratio around 0.50.

Prior to formulating testable structural hypotheses on the cointegrating vectors, two preliminary steps are taken. First, the impulse response functions derived from estimation and simulation of the *VAR* in (9), with the reduced rank restriction ($r = 2$) imposed, are examined. A simple Choleski decomposition of the residual covariance matrix is used to obtain orthogonal disturbances; the ordering chosen is: R^b, R^m, y, $m-p$, but the long-run responses of the variables are qualitatively robust to changes in the ordering. The estimated impulse response functions over a sixty-month horizon are shown in Figure 4 together with 95% confidence bounds: the four columns depict the responses of the four endogenous variables of the system to a shock in R^b, R^m, y and $m-p$ respectively. The two long-run features of the data highlighted above are confirmed: (*i*) the long-run responses of money balances and income to all disturbances are in the same direction and have similar magnitude; (*ii*) the response of the Treasury bill rate to all shocks is almost twice as large as that of the net yield on M2 (R^m). Moreover, when the interest rates show permanent long-run reactions to some disturbances (e.g. to a shock in R^b, in the first column of Figure 4), neither money balances nor income do seem affected in a quantitatively important way. Similarly, when the disturbances have permanent effects on $m-p$ and y (as in the last two columns), the two interest rates do not display in the long-run a statistically significant response.



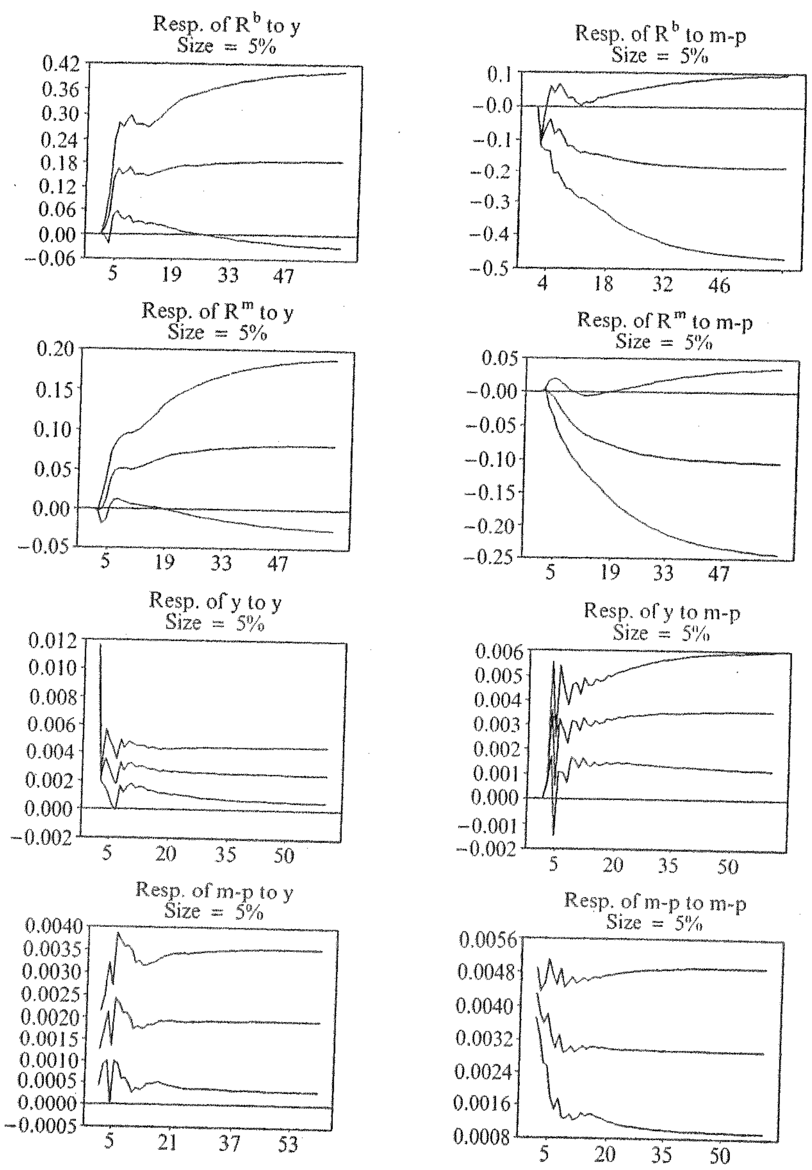


Figure 4: Impulse response functions from the cointegrated *VAR*

Notes: The impulse response functions are derived from simulation of the *VAR* system in equation (9) over a sixty-month horizon. Orthogonal disturbances are obtained by means of a Choleski decomposition of the *VAR* residual matrix with the variables ordered as: R^b, R^m, y, $m-p$. The four plots in each column show the responses of one endogenous variable to a one-standard deviation disturbance in R^b, R^m, y, $m-p$ respectively, with 95% confidence bounds. The standard deviations of the disturbances are: R^b: 0.37; R^m: 0.09; y: 0.010; $m-p$: 0.004.

Therefore, as a second preliminary step, we investigate whether the long-run relations between pairs of variables, seemingly strong features of the data, may form *by themselves* valid cointegrating vectors. This is done by applying cointegration analysis separately to two sub-systems of variables: $(m-p, y)$ and (R^m, R^b). In each sub-system, beside the inflation rate, the dummy and exogenous variables relevant to modelling the endogenous variables (see Table 1) are included in estimation as additional stationary regressors. Moreover, in the $(m-p, y)-(R^m, R^b)$ – system, three lags of ΔR^m and $\Delta R^b - \Delta(m-p)$ and Δy – are added in order to allow for more general short-run dynamics, since the omission of important short-run effects may in principle invalidate the estimation of the long-run properties of sub-systems of variables (Johansen and Juselius [1992]). The results[10] show that both pair of variables are cointegrated. In the money-expenditure system the normalized coefficient on y is 0.73, somewhat lower than those estimated from the complete *VAR*. In the interest rates system the relative magnitude of the estimated coefficients is again around 0.50. The issue of the proper specification of the deterministic component is addressed in the context of the two sub-systems by including a linear time trend in the money-income system and a constant in the interest rate system. When formally tested both deterministic terms are found not statistically significant, justifying their exclusion from the specification of the cointegrating space in Table 2.

These findings suggest that the original estimates of the cointegrating vectors in β from the complete system may then be linear combinations of two underlying distinct long-run relations, one between real money balances and total final expenditure and the other linking the interest rates on Treasury Bills and on M2. The latter relation may capture banks' behavior in setting the interest rate on deposits with reference to the bill rate, whilst the former may be justified on the basis of transaction money demand theory. In fact, models of the purely transactive motive for money holding of the Baumol-Tobin variety yield a well-determined negative relation between the interest rate on alternative assets (or the interest rate differential) and money balances. In the original contributions by Baumol (1952) and Tobin (1956) the interest rate negatively affects money demand through changes in the frequency of withdrawals of funds from interest-bearing assets, with an exogenously fixed pattern of spending between withdrawals. However, this basic model may be generalized, as in Romer (1986), by allowing utility-maximizing consumers to choose simultaneously the number and timing of bond

[10] An *Appendix*, available upon request, reports the detailed results of the sub-system analysis.

conversions into money (necessary for transactions purposes), the amount of each conversion and the pattern of consumption between conversions. In this extended framework the interest rate affects also the money holding pattern between conversions and the size of conversions. The latter (*wealth*) effect positively links in the long-run money holdings to the (alternative) interest rate and may offset the other negative effects, working through changes in the ratio of average money holdings between conversions to the initial amount transformed from bonds into money and through changes in the frequency of conversions, along the traditional Baumol-Tobin lines[11]. Therefore, more general versions of the traditional model of the transactions demand for money, in the presence of a sufficiently strong wealth effect, may generate small negative (or even positive) values of the interest rate elasticity of money demand. In this perspective, and on the basis of our preliminary result of a stationary relation involving money balances and expenditure in a bivariate system, we impose the restriction of an empirically negligible (formally zero) long-run interest rate effect on real money holdings in the complete four-variable *VAR*. Furthermore, we note that thinking of M2 money holdings as mainly motivated by transaction purposes is in accordance with some recent empirical evidence on Italian money demand behaviour, obtained with more conventional methods. In the context of a single equation analysis, Angelini *et al.* (1994) reach the conclusion that in the '80s M2 has fulfilled mainly the role of transaction medium, whereas until the late '70s money balances served also as a store of value, due to the limited set of alternative financial assets and the lack of liquid secondary markets for the existing instruments[12]. The process of financial innovation occurred in the late '70s and early '80s determined a widening of the range of financial assets available to investors (mainly through the introduction of Treasury's floating rate certificates (*CCT*)) and the development of a more liquid and efficient secondary market for the already used Treasury bills (*BOT*). This resulted in sizeable reallocations of private sector portfolios away from money. Angelini *et al.* (1994) empirically characterize this process as a gradual shift from (a measure of) financial wealth to final expenditure as the relevant scale variable in the estimated

[11] In principle, even in the original Baumol-Tobin model (first extended to allow for wealth effects by Johnson [1970]) money demand can be interest-inelastic, with a unitary income elasticity, if the frequency of income receipts is sufficiently high that agents never find it convenient to put a portion of their income into interest-bearing assets to be subsequently liquidated.

[12] As suggested by a referee, the significant long-run interest-rate effect on broad money demand found by Muscatelli and Spinelli (1996) using data from 1860 to 1990 is in accord with the above view on the changing role of M2 balances.

equations for M2. Moreover, Terlizzese (1994), in the context of a small-scale version of the quarterly econometric model used by the Bank of Italy for policy analysis, adopts a specification for real M2 demand in the post-1983 period based exclusively on the transactions motive and with the interest rate on Treasury bills (the only rate in the equation) affecting only the short-run dynamics of money balances, with no long-run effect[13].

We now provide a formal evaluation of the long-run structural hypotheses on money demand and interest rate behaviour formulated above. Our testing procedure involves three related steps and makes use of the likelihood ratio tests described and applied by Johansen and Juselius (1992, 1994).

A) First we test the hypothesis that *one* of the cointegrating vectors has a given form, leaving the other vector unrestricted. According to the previous discussion, two specific hypotheses are tested:

A1) The cointegration space spanned by the columns of β contains a vector of the form $(a, b, 0, 0)$, for some a and b to be estimated. This amounts to testing whether a linear combination of money balances and expenditure *alone* may be considered as a valid cointegrating relation in the complete system, leaving the second vector totally unrestricted;

A2) The cointegration space spanned by β contains a vector of the form $(0, 0, c, d)$, for some c and d to be estimated. This tests the existence of a valid cointegrating relation between the two interest rates, with no role for money balances or expenditure.

As shown by Johansen and Juselius (1992) a likelihood ratio test statistic for the above hypotheses can be constructed from the estimated eigenvalues under the restricted and unrestricted models. The test results (Table 3, panel A) show that neither hypothesis may be rejected. The relation between real money balances and expenditure is a valid cointegrating vector (A1); after normalization on $m-p$, the estimate of the coefficient on y is 0.88. The unrestricted vector displays the pattern of interest rate coefficients observed in the original estimates of β, with the coefficients on $m-p$ and y very close to zero. When the zero restrictions on money balances and expenditure in one cointegrating vector are imposed (A2) the estimated coefficient on R^b (normalized on R^m) is 0.495. The unrestricted vector has coefficients close to zero on the two interest

[13] In recent years, M2 growth is attributable mainly to its less liquid components, namely certificates of deposit with maturity longer than eighteen months (Bank of Italy [1993]). The high degree of substitutability of these assets with other financial instruments not included in the M2 definition does cast some doubt on the possibility of extending beyond 1991 the interpretation of M2 holdings as an essentially transactions-motivated.

rates and a relation between the coefficients on $m-p$ and y not very different from that found under A1, although the estimate of the coefficient on y (1.06, after normalization on $m-p$) is somewhat higher than that obtained under A1.

B) Given the above findings, we proceed to test the hypothesis that each one of the detected relations between pairs of variables ($m-p$ and y on the one hand, and R^m and R^b on the other) enter *all* cointegrating vectors. Hence, the following two hypotheses are tested:

B1) In both cointegrating vectors the coefficients on $m-p$ and y are proportional to $(1, -a)$, with $a = 0.88$, the estimated coefficient under A1 above, so that the cointegrating relations have the form $(z, -0.88z, *, *)$;

B2) In both cointegrating vectors the coefficients on R^m and R^b are

Table 3

Structural Restrictions on Long-run Relations
(* denotes imposed parameter restrictions)

A. Restrictions on *one* cointegrating vector.

A1. Two zero restrictions for coefficients on R^m and R^b

Restricted estimated cointegrated vectors
(restricted vector normalized on $m-p$, unrestricted vector normalized on R^m)

$m-p$	1.943 (−1)	0.481 (−0.208)
y	−1.710 (0.880)	−0.358 (0.154)
R^m	0*	2.315 (−1)
R^b	0*	−1.085 (0.469)

LR test of restrictions: $\chi^2(1) = 0.010$ (p-value: 0.92)

A2. Two zero restrictions for coefficients on $m-p$ and y

Restricted estimated cointegrated vectors
(restricted vector normalized on R^m, unrestricted vector normalized on $m-p$)

$m-p$	0*	0.496 (−1)
y	0*	−0.526 (1.060)
R^m	−7.857 (−1)	−0.152 (0.306)
R^b	3.894 (0.495)	−0.010 (0.020)

LR test of restrictions: $\chi^2(1) = 0.554$ (p-value: 0.46)

(contd.)

(contd.)

B. Restrictions on *all* cointegrating vectors.

B1. Imposed restriction: (coeff. on y) $= -0.880$ (coeff. on $m-p$)

Restricted (standardized) estimated cointegrated vectors
(in parentheses coefficients normalized on R^m)

$m-p$	−1	−1
y	0.880*	0.880*
R^m	4.160 (−1)	−1.517 (−1)
R^b	−2.099 (0.505)	0.774 (0.510)

LR test of restrictions: $\chi^2(2) = 4.60$ (p-value: 0.10)

B2. Imposed restriction: (coeff. on R^b) $= -0.495$ (coeff. on R^m)

Restricted (standardized) estimated cointegrated vectors
(in parentheses coefficients normalized on $m-p$)

$m-p$	0.161 (−1)	−0.641 (−1)
y	−0.161 (0.999)	0.543 (0.847)
R^m	−1	−1
R^b	0.495*	0.495*

LR test of restrictions: $\chi^2(2) = 1.32$ (p-value: 0.52)

C. *Fixed cointegrating vectors*

Imposed matrix of coefficients:

$m-p$	−1	0
y	0.880	0
R^m	0	−1
R^b	0	0.495

LR test of restrictions: $\chi^2(4) = 4.78$ (p-value: 0.31)

proportional to $(1, -b)$, with $b = 0.495$, the value found under A2, so that the cointegrating relations have the form $(*, *, z, -0.495z)$. The values of the appropriate test statistics show that neither hypothesis is rejected and the estimates of the unrestricted coefficients (on R^m and R^b under B1 and on $m-p$ and y under B2) confirm the patterns previously detected.

C) Finally, we assume that *both cointegrating vectors* are *known*: one is proportional to $(-1, a, 0, 0)$ with $a = 0.88$ and the other is proportional to $(0, 0, -1, b)$ with $b = 0.495$. The appropriate likelihood ratio statistic gives a value of 4.78 with a corresponding probability value of 0.31.

Overall, the above results give some support to the view that the two valid cointegrating vectors involving the four variables under study are of the form given under (C): one describes a long-run relation between real money balances and expenditure, interpretable as a simple transactions demand for money with a point estimate for the expenditure elasticity of 0.88 and interest rate elasticities not significantly different from zero, the other essentially capturing the long-run tendency of interest rates on deposits to reflect movements in market rates with a coefficient of 0.5. These relations are used to construct the following disequilibrium (error-correction) terms, to be included in the system analysis of the next section:

$$\begin{aligned} ECMM &= (m-p) - 0.880y \\ ECMR &= R^m - 0.495R^b \end{aligned} \tag{11}$$

$ECMM$ and $ECMR$ measure the (short-run) deviations of money balances and R^m from their long-run equilibrium level as determined respectively by expenditure and by R^b [14].

Before proceeding further, cointegration analysis is used to settle two modelling issues raised by the chosen VAR specification in (9). First, we test long-run price-level homogeneity in the context of a three-variable cointegrated system including separately nominal money (m), the price level (p) and expenditure (y) [15]. Johansen's estimation procedure applied to this system reveals, as expected, the presence of only one cointegrating vector. The hypothesis of a unit coefficient on the price level (once normalized on m) is then tested by means of a likelihood ratio test of the kind used for hypotheses A1 and A2 above. The resulting value of the test statistic is 0.50, with a corresponding probability value of 0.48. We therefore conclude that price level homogeneity is not rejected and, consequently, our choice to specify the monetary aggregate in real terms is consistent with the long-run

14 Alternative long-run hypotheses were also tested. Two results are worth mentioning: *i*) the hypothesis that the coefficient on y in the money-expenditure cointegrating vector is 1 (a *velocity* restriction) is not rejected, whereas *ii*) the hypothesis that the interest rate differential is a stationary relation is strongly rejected. The result under *i*) suggests that the value of the expenditure elasticity is not very precisely determined; in the following analysis we use the value obtained under A1 above (0.88). However, the conclusions of the next section are unchanged when a unitary coefficient on y is imposed in the $ECMM$ term. Furthermore, the presence of a linear time trend in the cointegrating space has been tested in the whole system obtaining a value of 7.9 for the $\chi^2(4)$ LR statistic (p-value: 0.10) when the absence of the trend is imposed onto the β matrix together with the exclusion restrictions on the two cointegrating vectors tested under C above.

15 Three lags of ΔR^m and ΔR^b are included in the estimated system as additional stationary regressors to allow for more general short-run dynamics.

properties of the data. The second issue concerns the stationarity of the inflation rate detected by the *ADF* test. To provide an additional test of this property we apply Johansen's procedure to an extended *VAR* system, with Δp included as an additional *endogenous* variable. Now, three valid cointegrating vectors are found, one more than in the four-variable system: this is consistent with an *I(0)* variable being included in a system of *I(1)* series. A formal test does not reject the hypothesis that Δp is the only variable entering one vector and is excluded from the other two cointegrating relationships (the associated p-value is 0.11). We interpret this result as further evidence of the *I(0)* nature of Δp, supporting our choice of omitting it from the long-run determinants of money demand.

2.3. *From the Cointegrated VAR to a Simultaneous Model*

The previous analysis has reached two main conclusions: (*i*) there is evidence of two long-run relations involving the endogenous variables of the system; (*ii*) the data do not reject simple structural hypotheses, suggested by the long-run properties of sub-systems of the variables. We therefore have an alternative to the single-equation procedure of taking the original estimates of the first cointegrating vector as a valid long-run money demand function and including the derived error-correction term in a dynamic equation for real money balances. In so doing, the existence of a second long-run relation among the variables (or maybe a subset there of) is neglected and information potentially contained in other equations of a multivariate system is ignored. On the contrary, we adopt a system approach and proceed to model the short-run adjustment of all endogenous variables towards their equilibrium relations, allowing for contemporaneous interactions between money, expenditure and interest rates. If the evidence on the long-run can be validly read as we did in the previous subsection, the dynamic adjustment to equilibrium must be consistent with the economic interpretation given to the long-run cointegrating relations. In particular, the disequilibrium (*ECM*) terms in (11) should determine a plausible pattern of error-correcting responses of the endogenous variables. This does not necessarily imply that each *ECM* term, constructed from a particular cointegrating vector, must enter only (some of) the equations corresponding to the variables belonging to that vector. In fact, deviations from the equilibrium path involving a subset of variables may have important short-run effects on the dynamics of other variables not included in the long-run equilibrium relation (Chow [1993], Konishi, Ramey and Granger [1993]). In what follows we formulate some hypotheses on the dynamic, short-run adjustment pattern of the variables, consistent with the interpretation of the long-run cointegrating vectors previously tested.

As a prerequisite for valid testing, we estimate the four-variable dynamic system in (9) with the two error-correction terms (lagged one period) in (11) replacing the unrestricted lagged levels of the endogenous variables. The short-run dynamics are left completely unrestricted. For this system, a *restricted cointegrated VAR*, to provide a suitable framework for the subsequent empirical analysis, it is necessary that the equation residuals are normally distributed innovation processes and the conditioning variables are *weakly* exogenous for the parameters of interest (Engle, Hendry and Richard [1983], Engle and Hendry [1993]).

Among the conditioning variables included in the system, exogeneity problems potentially arise only for the inflation rate, since it can be plausibly assumed that there is no contemporaneous (within-month) feedback from activity and real money balances to monetary policy actions, captured by changes of the discount and repo rates (in fact, aggregate statistical information on the behaviour of the economy is available to monetary authorities only with at least a month's delay). We test for the weak exogeneity of Δp for the parameters describing the short-run dynamics of the system following Engle and Hendry (1993). Our aim is to test that there is no loss of information in conditioning the system on the inflation rate, so avoiding the joint modelling of an additional variable. Formally, this is done by formulating a time-series model for Δp, from which estimates of the parameters (mean and variance) of the marginal distribution are derived. For Δp to be weakly exogenous, the parameters of its marginal distribution must not enter the conditional system. The estimated marginal model for the inflation rate contains three lags of Δp and of each of the four endogenous variables in the system ($\Delta(m-p)$, Δy, ΔR^m and ΔR^b) and all the dummy and exogenous variables included in the system. The fitted values and the squared fitted values so obtained as proxies for the mean and variance of the distribution of Δp, are added to the *VAR* estimated above and tested for statistical significance. In all four equations the added terms are not significant both individually and jointly, supporting the conclusion of weak exogeneity of the inflation rate. Furthermore, in order to validate the forecast analysis and tests conducted on the system in subsection 2.1., *strong* exogeneity of the inflation rate is needed. Therefore, tests of Granger-causality from the endogenous variables in the system to Δp are carried out using three lags of each variable. The results show that none of the variables Granger cause the inflation rate, supporting the strong exogeneity of Δp. The same conclusion holds also for the policy rates included in the system.

A simplification of the general dynamics of the restricted *VAR* is performed by eliminating those regressors ($\Delta(m-p)_{t-2}$, ΔR^b_{t-1}, and

ΔR^b_{t-2}) having non-significant (system) F-test statistics and entering each individual equation with non-significant coefficients. The resulting system – a *parsimonious VAR* (*PVAR*) in Clements and Mizon (1991) terminology – is then estimated and F-tests for the statistical significance of the retained regressors are carried out and reported in Table 4, panel A. As can been seen from the high values of the corresponding F statistics, an important part of the explanatory power lies with the error-corection terms. All other regressors now display acceptably high levels of statistical significance, with perhaps the only exception of Δy_{t-2} and ΔR^m_{t-2} (the p-values are 0.31 and 0.36 respectively): these are nevertheless retained in the parsimonious version of the system, being important explanatory variables in at least one equation, as will be confirmed by the simultaneous model estimation. The *PVAR* residuals do not display deviations from normality and only in the expenditure equation (but not in the system as a whole) is some residual serial correlation detected (Table 4, panel B). The test statistic for the exclusion restrictions imposed in the *PVAR* provides formal support for the system reduction. Stability of the system is assessed by a recursive break-point Chow test, showing no evidence of structural breaks.

The *PVAR* can therefore be considered a suitable statistical framework whereby tests of simultaneous structural models may be validly carried out[16]. The general formulation of such a model is the following:

$$D_0 \Delta x_t = \Gamma \begin{pmatrix} ECMM_{t-1} \\ ECMR_{t-1} \end{pmatrix} + \sum_{i=1}^{3} D_i \Delta x_{t-i} + G_1 d_t + G_2 \Delta w_t + c + \eta_t \tag{12}$$

where $\Delta \mathbf{x}_t = \{\Delta (m-p)_t, \Delta y_t, \Delta R^m_t, \Delta R^b_t\}'$, $\mathbf{d}_t$ and $\Delta \mathbf{w}_t$ are vectors of dummy and exogenous variables respectively, and the error-correction terms *ECMM* and *ECMR* are defined in (11). The coefficients of the 4×2 matrix Γ capture the reaction of each endogenous variable to deviations from the two long-run equilibrium relations specified in section 2.2. $\mathbf{D}_0$

[16] As noted by Sims (1991), in the econometric literature, the term *structural* is used to denote models explicitly built on economic theories of optimizing behavior, with the estimated parameters directly related to characteristics of agents' tastes and technology. Moreover, in the context of *VAR* modelling, a *structural* model offers a behavioral interpretation to the various sources of stochastic disturbances in a multivariate *VAR*. We refer to the model below as *structural* in the (more limited) sense of embodying some behavioral hypotheses on the long-run equilibrium and some restrictions on the dynamic adjustment of the system towards such equilibrium, also based on a behavioral interpretation.

Table 4

A. *F-tests (and p-values) on retained regressors in the parsimonious VAR system: F (4,85)*

$\Delta (m-p)_{t-1}$	$\Delta (m-p)_{t-3}$	Δy_{t-1}	Δy_{t-2}	Δy_{t-3}	ΔR^m_{t-1}
2.13 (0.084)	2.76 (0.033)	3.98 (0.005)	1.21 (0.314)	2.60 (0.042)	3.15 (0.018)
ΔR^m_{t-2}	ΔR^m_{t-3}	ΔR^b_{t-3}	$ECMM_{t-1}$	$ECMR_{t-1}$	Δp_t
1.10 (0.360)	4.15 (0.004)	2.19 (0.076)	10.98 (0.000)	13.22 (0.000)	6.36 (0.000)
$\Delta DISC_t$	$\Delta REPR_t$	$\Delta DISCN_{t-1}$	$\Delta REPRN_{t-1}$	DUS_t	$DU878_t$
8.39 (0.000)	4.04 (0.005)	18.48 (0.000)	1.74 (0.148)	7.20 (0.000)	3.98 (0.005)
	$DU8310_t$	$DURM3_t$	$DU877_t$	$DU8967_t$	
	12.96 (0.000)	9.47 (0.000)	4.31 (0.003)	9.00 (0.000)	

B. *Residual mis-specification tests on the parsimonious VAR systems (p-values in parentheses):*

Statistic	Parsimonious VAR				
	Equation				*VAR*
	$\Delta (m-p)$	Δy	ΔR^m	ΔR^b	
σ	0.391	1.026	0.056	0.240	
AR 12 *F* (12,73)	1.07 (0.39)	2.20 (0.02)	1.33 (0.22)	1.18 (0.31)	
Normality χ^2 (2)	4.60 (0.10)	1.55 (0.46)	1.04 (0.59)	1.12 (0.57)	
Heterosc. *F* (32,52)	0.35 (1.00)	0.43 (0.99)	1.19 (0.28)	0.80 (0.74)	
ARCH (7) *F* (7,71)	0.41 (0.89)	0.19 (0.99)	0.61 (0.75)	0.57 (0.78)	
AR (12) *F* (192,138)					1.07 (0.34)
Normality χ^2 (8)					6.74 (0.57)
Heterosc. *F* (320,438)					0.76 (0.99)

Notes: In the last column of panel B mis-specification tests are conducted at the whole system level for twelve-order serial correlation, normality and heteroscedasticity. Dummy and exogenous variables in the *PVAR* are defined in notes **B** and **C** to Table 1.

Table 5

Simultaneous Model (FIML Estimation)

A. *Coefficient estimates (standard errors)*

Variable:	Equation for: $\Delta(m-p)_t$	Δy_t	ΔR^m_t	ΔR^b_t
Δy_t				0.056 (0.031)
ΔR^m_t	1.314 (0.512)			
ΔR^b_t		−0.554 (0.345)		
$ECMM_{t-1}$	−0.168 (0.035)	0.312 (0.075)		−0.057 (0.020)
$ECMR_{t-1}$	0.893 (0.139)		−0.082 (0.013)	
$\Delta(m-p)_{t-1}$		−0.414 (0.188)		−0.091 (0.048)
$\Delta(m-p)_{t-3}$		−0.314 (0.194)	−0.016 (0.010)	0.151 (0.050)
Δy_{t-1}	−0.073 (0.030)	−0.494 (0.086)		
Δy_{t-2}		−0.192 (0.074)		
Δy_{t-3}	−0.602 (0.024)			
ΔR^m_{t-1}	0.620 (0.373)		0.165 (0.052)	
ΔR^m_{t-2}			−0.076 (0.052)	−0.517 (0.221)
ΔR^m_{t-3}		0.380 (0.080)		0.379 (0.242)
ΔR^b_{t-3}		−0.586 (0.259)	−0.037 (0.015)	
Δp_t	−1.088 (0.214)	−1.526 (0.483)	0.093 (0.028)	
$\Delta DISC_t$				0.561 (0.082)
$\Delta REPR_t$				0.109 (0.027)
$\Delta DISCN_{t-1}$			0.210 (0.022)	
$\Delta REPRN_{t-1}$			0.033 (0.009)	
DUS_t	1.579 (0.305)			
$DU878_t$		−4.780 (0.040)		
$DU8310_t$			−0.426 (0.057)	
$DURM3_t$			−0.223 (0.033)	
$DU877_t$				1.176 (0.260)
$DU8967_t$				−1.311 (0.182)
σ	0.441	1.077	0.058	0.259

(contd.)

(contd.)

B. *Residual mis-specification tests on the simultaneous model*
(p-values in parentheses)

Statistic	Equation: $\Delta(m-p)$	Δy	ΔR^m	ΔR^b	*Model*
Ser. Cor. $\chi^2(12)$	11.38 (0.50)	19.73 (0.08)	17.69 (0.13)	12.12 (0.44)	
Normality $\chi^2(2)$	0.82 (0.66)	1.10 (0.58)	2.17 (0.34)(0.27)	2.63	
Heterosc. $F(39,45)$	0.34 (1.00)	0.41 (1.00)	1.02 (0.47)	0.70 (0.87)	
ARCH (7) $F(7,71)$	0.56 (0.78)	0.21 (0.98)	0.42 (0.88)	0.76 (0.62)	
AR (12) $F(192,190)$					0.96 (0.61)
Normality $\chi^2(8)$					6.68 (0.57)
Heterosc. $F(390,504)$					0.89 (0.89)

LR test of overidentifying restrictions: $\chi^2(50) = 46.9$ (0.60)

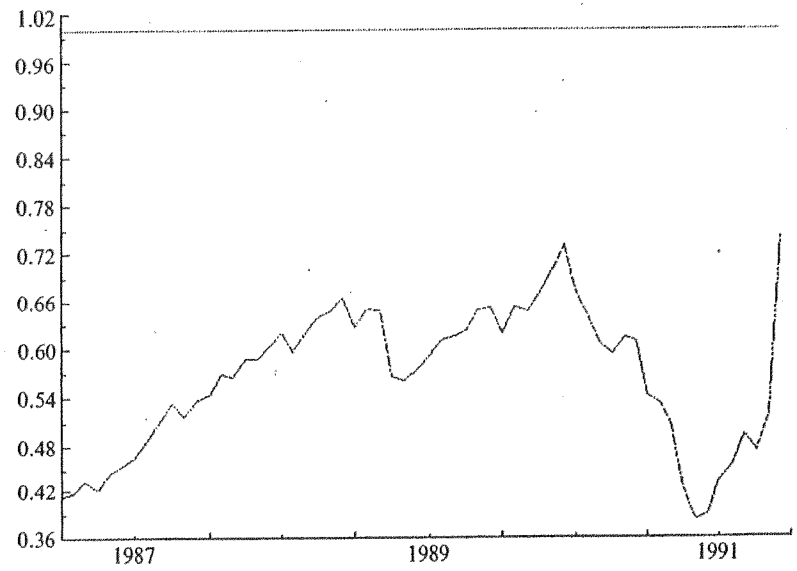


Figure 5: Break-point Chow stability test from recursive simultaneous model estimation: 1987-1991 (1.0 = 5% crit. value of the test)

the recursive estimates of the simultaneous model (Figure 5) confirm the absence of structural breaks over the 1987-1991 period.

In order to compare the results obtained from the multivariate approach employed here with those yielded by a more conventional single-equation analysis, we estimated an equation for money demand starting from an unrestricted general dynamic model with four lags of all variables involved (m, p, y, R^m, R^b). A general-to-simple modelling strategy is then followed in order to restrict the dynamics of the equation, using exclusion restrictions and reparameterizations of the original regressors and testing each successive step in the reduction process. The final result is the following equation (estimated by OLS over the 1983(1)-1991(12) sample period):

$$
\begin{aligned}
\Delta (m-p)_t = \; & \underset{(0.046)}{-0.201}\,(m-p)_{t-1} \underset{(0.031)}{+0.164}\, y_{t-1} \underset{(0.201)}{+0.715}\, R^m_{t-2} \underset{(0.088)}{-0.380}\, R^b_{t-2} \underset{(0.031)}{+0.095}\, \Delta y_t \\
& \underset{(0.026)}{-0.055}\, \Delta y_{t-3} \underset{(0.444)}{+1.478}\, \Delta R^m_{t-1} \underset{(0.303)}{-0.775}\, \Delta p_t \underset{(0.331)}{+1.575}\, DUS_t \underset{(19.52)}{-5.724}
\end{aligned}
$$

$$R^2 = 0.370 \qquad \sigma = 0.437$$

Diagnostic Tests (p-value)

Serial correlation	$F(12,86) = 1.25$ (0.26)	*Funct. form*	$F(1,97) = 0.55$ (0.46)
Normality	$\chi^2(2) = 2.86$ (0.24)	*Heterosc.*	$F(1,106) = 0.39$ (0.53)
Predictive failure	$F(12,98) = 2.235$ (0.016)		

The diagnostic tests reported show that the only problem affecting this single-equation specification of money demand is, as expected, predictive failure over the 1992(1)-1992(12) period [17].

The coefficients on the regressors in levels yield the following long-run solution:

$$m - p = \underset{(0.088)}{0.814}\, y + \underset{(1.374)}{3.556}\, R^m - \underset{(0.604)}{1.890}\, R^b$$

This linear combination of the four variables analysed is conventionally interpreted as a long-run money demand function with plausible elasticities. Here a long-run response of money balances to both interest rates is detected, in addition to an elasticity to expenditure lower than unity. In fact, in the light of our system analysis, this long-run solution of the model may be viewed as the particular linear combination of the two underlying cointegrating relations among the variables which is supported by the data. In fact, the value of the expenditure elasticity is not very different from the one obtained in the multivariate cointegration analysis and also the ratio of the two interest rate coefficients (0.53) reproduces almost exactly the coefficient linking the two rates in the long-run.

[17] An $F(15,83)$ test of the 15 parameter restrictions of the final specification against the general unrestricted model yields a value of 0.90. When estimation is performed by IV methods, instrumenting Δy_t with lags of itself, of $\Delta(m-p)$, ΔR^m and ΔR^b, and with the dummy variable $DU878$, the results are unchanged.

The multivariate analysis offers an alternative interpretation of the data which, though yielding a specification of money demand behaviour which is observationally equivalent to the conventional single-equation money demand function, has the advantage of using information from *all* equations in the system, accounting for the multiplicity of long-run relations. Moreover, the consistency of the short-run adjustment process for all variables with the economic interpretation of the equilibrium path of the system can lend support to the results obtained from estimation of the multivariate dynamic model.

Concluding Remarks

It is widely recognized that money balances, expenditure and interest rates may be linked by multiple long-run relations. This possibility makes it difficult to give a structural interpretation to the results from single-equation studies of money demand. A multivariate framework is needed to detect such relations and formally test economic hypotheses on the long-run features of the data. Once a (non-rejected) structural interpretation of the equilibrium relations in the system is obtained, a complete simultaneous dynamic model for all variables may be specified and evaluated. The short-run adjustment dynamics of the system must be consistent with the proposed economic interpretation of the long-run equilibrium.

This methodology, combining cointegration analysis with more traditional structural modelling, is applied to Italian data for the eighties and early nineties. The results show that the short-run time-series behavior of money balances, expenditure and interest rates may be described as adjusting towards two equilibrium relations, one between real money holdings and expenditure (interpretable as a simple transactions demand for money) and the other linking the yields on money and on Treasury bills. The dynamic adjustment of the variables is readily interpretable: money holdings and expenditure react in an error-correcting fashion to deviations from the money-expenditure equilibrium path, whereas deviations from the long-run relation linking the yield on money to the bill rate determine an equilibrating response of the interest rate on money and also affects money holdings dynamics in the short-run. The pattern of dynamic responses of the variables to deviations from the system's long-run equilibrium is viewed as supporting the economic interpretation of the multiple cointegrating relations.

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