STRUCTURAL STABILITY, SHORT-RUN DYNAMICS AND LONG-RUN SOLUTIONS IN FEEDBACK MODELS. THE CASE OF THE MONEY DEMAND FUNCTION FOR ITALY: 1964-1986 *

Introduction

Empirical models based on feedback rules, whereby agents react to observed (lagged and current) variables when deciding current values of their choice variables, have been widely used in applied econometrics. One popular justification for the adoption of this class of models, put forward, for example, by Hendry (1988), is that they represent simplified rules-of-thumb which agents may follow in complex environments. On practical grounds, several strategies have been formulated in order to obtain satisfactory empirical specifications of such behaviour. They are all based on the recognition of the existence of a long-run, equilibrium relation between the decision variable to be modelled and its determinants. However, adjustment costs and other (perhaps informational) imperfections prevent such a relation from being satisfied at every moment in time and give rise to a maybe complex short-run dynamics around the long-run equilibrium.

A well defined econometric strategy for feedback models should therefore be capable of capturing both the equilibrium condition and the form of the short-run dynamics. Two such specification strategies are investigated in this paper. The first, mainly due to the work of D. Hendry, aims at a simultaneous specification of the long-run relation and the short-run dynamics, whereas the second, recently developed by Engle and Granger (1987), reaches the final specification through a two-step procedure, whereby long-run and short-run elements are sequentially investigated.

As some recent literature — to be discussed below — has pointed out,

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the whole class of feedback models, ignoring the potential role of expectations in determining current behaviour, may suffer from serious drawbacks. These may concern both the economic interpretation of the estimates of long-run equilibrium parameters and the possibility of confidently using the estimated equations for policy evaluation (a point forcefully made by Lucas (1976)).

In the present paper, several specifications of a feedback model of the demand for money in Italy over the period 1964-1986 are presented and contrasted. The main focus of our analysis is on some unsatisfactory features of these specifications which seem to suggest the need for the inclusion of expectations variables in the equation and the explicit modelling of the expectations generating process adopted by economic agents. Particular emphasis is placed throughout on the performance of various feedback specifications under structural stability analysis and on the interpretation of the long-run equilibrium solutions obtained.

The choice of the demand for money function in order to evaluate some of the potential problems affecting purely feedback models seems particularly appropriate, since in the recent applied econometric literature several feedback specifications of money demand have been presented and considered as satisfactory representations of the underlying behaviour of agents. The best-known are perhaps those of Hendry (1985) for the United Kingdom and of Rose (1985) and Baba, Hendry and Starr (1988) for the United States. Furthermore, the potential relevance of expectations in the modelling of money demand has been recently investigated by Cuthbertson and Taylor (1987) and Cuthbertson (1988), who successfully constructed explicit forward-looking models of the demand for narrow money in the United Kingdom.

Even though the construction and estimation of a full-fledged expectations model goes beyond the scope of the present paper, our analysis of alternative feedback specifications suggests that more satisfactory results, particularly in terms of structural stability, could be obtained by explicitly considering the role of expectations in determining the demand for money.

We start in Section 1 by presenting the two widely adopted specification strategies for feedback models mentioned above and by briefly reviewing some of the problems stemming from the neglect of expectations. In Sections 2 and 3 these methodologies are applied to the demand for money (M2) in Italy and the performance of the final specifications is assessed by various criteria. In particular, a recursive stability analysis shows that all final equations are affected by several marked structural breaks. Section 4 contains some comments on a feedback specification recently proposed by Papi (1988), obtained using a modified version of the simultaneous methodology described in Section 2. Again, the same pattern of structural breaks occurs. Section 5 concludes the paper.

1. The econometric specifications of a feedback model

A feedback model can be defined as a marginalized and conditioned representation of the Data Generating Process (DGP) which includes only observed variables and is not based on the explicit modelling of expectations.

The basic principle in approaching the problem of the econometric specification of a feedback model is given by «congruency» with the data. According to Hendry and Richard (1983), a specification is congruent when the information contained in the data and omitted from the model is not relevant to the problem at hand.

To be a congruent representation of the data, a feedback model must capture both the long-run equilibrium relation between the variables under study and the shape of the short-run dynamics. In the recent literature, several methods have been proposed to achieve this result.

One common feature of all strategies is that the final specification is meant to be a «balanced» representation of the data, in the sense that the statistical properties of the dependent and explanatory variables included in the model must be internally consistent. In particular, in order to apply classical asymptotic results, stationary variables are needed. Since most economic variables are non-stationary, balanced relations between stationary variables can be achieved by appropriate differentiation of non-stationary variables or by considering cointegrating vectors, *i.e.* stationary linear combinations of non-stationary variables[1].

Specification strategies can be classified into two main groups, according to whether the long-run equilibrium and the short-run dynamics are modelled jointly or in successive steps.

1.1. *Simultaneous Specification of Long-Run Equilibrium and Short-Run Dynamics*

The simultaneous specification of the long-run equilibrium relation and the short-run dynamics fits well into the «general to specific» modelling strategy developed by D. Hendry in several papers (Hendry (1985), 1987)). The philosophy of this strategy is to derive the model from the DGP through several steps of reduction. The process of reduction involves a loss of information, which can or cannot be relevant to the scope of the researcher. The relevance of this loss of information can be assessed by testing procedures, designed to check if the model is a structurally stable representation of the DGP in which the error is a true innovation and the regressors satisfy the

[1] For a precise illustration of the concept of cointegration and a discussion of its statistical foundations see R.F. Engle and C.W.J. Granger (1987) and the surveys by C.W.J. Granger (1986), and J. Dolado and T. Jenkinson (1987).

exogeneity requirement for the relevant parameters which is appropriate for the proposed use of the model[2].

In practice, this procedure is implemented by starting from a general «baseline» model including long lags of both the dependent and the independent variables. The generality of the model reflects the belief that economic theory can only suggest which variables have to be related, but only the data can determine the precise dynamic relationship between them.

Note that the general model is already the outcome of some reduction: in fact it has been marginalized with respect to all the variables included in the DGP but omitted from it. Therefore, the diagnostic checking procedures designed to test the relevance of lost information have to be implemented also on the baseline model.

Once the congruency of the general model has been assessed by the diagnostic checking, further reductions can be implemented by imposing all the restrictions suggested by the data in the form of both exclusion restrictions and transformations on the level of the variable (differences or Error-Correction terms).

As an example, consider the problem of modelling a non-stationary $I(1)$ series y by a vector of forcing non-stationary $I(1)$ variables $\mathbf{x}$.

The general baseline model will have the following form

$$y_t = a + \sum_{i=i}^{M} \alpha_i y_{t-i} + \sum_{i=0}^{M} \boldsymbol{\beta}_i' \mathbf{x}_{t-i} + u_t \tag{1}$$

where a is a constant, the $\boldsymbol{\beta}_i$'s are 1 by n vectors of coefficients and u_t is a true innovation. A possible final outcome of the reduction process can be the following[3]

$$Dy_t = a + \boldsymbol{\beta}_0' \mathbf{Dx}_t + \delta (y_{t-1} - \mathbf{k}' \mathbf{x}_{t-1}) + \epsilon_t \tag{2}$$

where ϵ_t is a true innovation, and the following set of restrictions on the baseline model cannot be rejected

$$\alpha_i = 0 \text{ and } \boldsymbol{\beta}_i = \mathbf{0} \quad \text{for } i > 1, \quad \alpha_1 = 1 + \delta, \quad \boldsymbol{\beta}_1 = -(\boldsymbol{\beta}_0 + \delta \mathbf{k}) \tag{3}$$

If y and $\mathbf{x}$ are cointegrated with cointegrating vector $(1, -\mathbf{k})$, equation (2) is a balanced relation, involving only first differences of $I(1)$ variables and the (stationary) cointegrating relation.

As far as the economic interpretation is concerned, a necessary condi-

[2] For a detailed discussion of the main features of D.F. Hendry's methodology see GILBERT (1986) and, for a critical evaluation, see A. PAGAN (1987). On the concept of exogeneity, see R.F. ENGLE, D.F. HENDRY and J.F. RICHARD (1983).

[3] From now on, given a generic variable x_t, $Dx_t = x_t - x_{t-1}$.

tion to sustain the feedback interpretation of (2) is that $\delta < 0$. In this case, it can be argued that there is a long-run equilibrium relation given by the cointegrating vector ($y = \mathbf{k}'\mathbf{x}$) and agents, when deciding the value y_t, react to past deviations from equilibrium in such a way that the change in y tends to correct for past errors, being positive when the disequilibrium term is negative and vice versa. This interpretation justifies the widely used denomination of Error Correction Mechanism (ECM).

On the other hand, a positive value for δ would make the ECM interpretation not sustainable. In fact, $\delta > 0$ would amplify any past disequilibrium, preventing the model from convergence. Nevertheless, some economic foundations of such an outcome can be provided in the framework of dynamic intertemporal optimization proposed by Nickell (1985).

To illustrate this point, consider a standard intertemporal quadratic adjustment cost model (as developed, for example, in Sargent (1978)), in which economic agents are supposed to make a sequence of decisions y_t designed to chase a target variable y_t^* which evolves according to the relation

$$y_t^* = \mathbf{k}'\mathbf{x}_t + u_t \tag{4}$$

We consider (4) as a balanced relation in the sense that the target and the variables designed to chase it are of the same order of integration. Moreover, agents are rational in the choice of the vector $\mathbf{x}_t$: they choose instruments to track down the target so that they are not consistently wrong and make u_t stationary. In other words (4) is a cointegrating relation.

The cost function to be minimized, incorporating both costs of being out of equilibrium and costs of adjustment, is the following

$$\min_{y_{t+s}} E_t \sum_{s=0}^{\infty} \Phi^s [c(y_{t+s} - y_{t+s}^*)^2 + (y_{t+s} - y_{t+s-1})^2] \tag{5}$$

where c a positive constant measuring the relative importance of the two types of costs and Φ represents the constant discount factor ($0 < \Phi < 1$).

The first order condition of the maximization problem can be reparameterized as follows

$$Dy_t = (1/\Phi)Dy_{t-1} + (c/\Phi)(y_{t-1} - \mathbf{k}'\mathbf{x}_{t-1}) + (Dy_t - Dy_t^e) - (c/\Phi)u_{t-1} \tag{6}$$

By considering $(Dy_t - Dy_t^e)$ as an innovation we can see that equation (6) is a restricted version of (1). Namely, the restrictions are

$$\alpha_i = 0 \quad \text{for } i > 2, \quad \beta_0 = 0, \quad \beta_i = 0 \quad \text{for } i > 1,$$

$$\alpha_1 = 1 + (1/\Phi) + (c/\Phi), \quad \alpha_2 = -(1/\Phi), \quad \beta_1 = -(c/\Phi)\mathbf{k}.$$

Notice that equation (6) implies an «Error Amplification» Mechanism,

(c/Φ) being positive: a level of y_{t-1} greater than its equilibrium level has a positive impact on the rate of growth of y. Equation (6) is not generated by agents correcting for past errors but is the outcome of their rational forward-looking behaviour, in the presence of adjustment costs.

This observation emphasizes the importance of solving the problem of observational equivalence between structural feedback models and feedback specifications which are reduced forms of structural feedforward models. Although in the above specific example the sign of the coefficient on the combination of the lagged levels provides a clear identifying restriction, in models with a more complex dynamics, it is perfectly possible to obtain a negative coefficient on the level term within a feedforward theoretical framework, as shown in Nickell (1985) and Dolado (1987).

If the feedback model is to be interpreted as the reduced form of a feedforward model, then the estimated coefficients are convolutions of «deep» behavioural parameters and expectations parameters. This fact gives rise to two major problems.

Firstly, as Kelly (1985) pointed out, the long-run elasticities derived from the feedback model are different from the true behavioural elasticities because of the presence of the expectational parameters.

Secondly, the feedback model is subject to the Lucas (1976) critique and no exercise of econometric policy evaluation is feasible.

However, Hendry and Neale (1988) have recently shown that, if the ECM term constitutes a cointegrating relation, the parameters in the cointegrating vector define a stationary linear combination of non-stationary variables and are not affected by whether observed or expected variables are included in the underlying structural relation. In fact, under rational expectations, actual and expected variables differ only by a stationary ($I(0)$) expectational error which by its nature does not affect the estimated cointegrating relation.

On the other hand, it is in general possible to discriminate between structural feedback model and feedback model which are reduced forms of feedforward structural models by using structural stability and encompassing tests. Hendry (1988) provides an illustration of such possibility and uses recursive structural stability tests to show that a feedback model for the United Kingdom money demand cannot be interpreted as a reduced form from a feedforward model of the kind proposed by Cuthbertson (1988) because the feedback specification does not show any sign of structural instability when the parameters in the expectations' generating processes show remarkable structural breaks.

1.2. *The Two-Step Procedure*

In the two-step procedure, the long-run equilibrium relation and the short-run dynamics are modelled sequentially. In the first step, the long-run equilibrium relation between the variables is estimated by means of a cointegrating regression. The residuals from such regression are then used as an ECM term around which the short-run dynamics are modelled in the second step.

This procedure, proposed by Engle and Granger (1987), exploits the property of cointegration implying a causal nexus between the cointegrated variables. This causal nexus represents a long-run equilibrium relation which can be used as a starting point in specifying a valid representation of the DGP.

However, a cointegrating vector is a static representation of the theory and we have to consider further developments in order to represent short-run dynamic fluctuations.

The basis for such a development is provided by the Granger-Engle Representation Theorem. One of the statement of the theorem is that if some series form a cointegrating vector then an Error Correction Mechanism representation is allowed.

To show the importance of this statement, consider the following representation of two autoregressive series x_t and y_t, cointegrated of order (1,1), with x_t weakly exogenous for the estimation of the cointegrating vector

$$y_t - \alpha_{12} x_t = u_{1t} \tag{7}$$

$$x_t = \beta_{21}(L) x_{t-1} + \beta_{22}(L) y_{t-1} + u_{2t} \tag{8}$$

$$\begin{aligned} u_{1t} &= \pi u_{1t-1} + \epsilon_{1t} \\ u_{2t} &= \quad u_{2t-1} + \varepsilon_{2t} \end{aligned} \quad \varepsilon_t = \begin{bmatrix} \epsilon_{1t} \\ \epsilon_{2t} \end{bmatrix} \sim N\left[\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} \sigma_{11} & 0 \\ 0 & \sigma_{22} \end{pmatrix}\right]$$

Cointegration implies that $|\pi| < 1$.

By differencing we obtain

$$\mathbf{A}\begin{bmatrix} Dy_t \\ Dx_t \end{bmatrix} = \mathbf{B}(L)[y_{t-1}, x_{t-1}, Dx_{t-1}, Dy_{t-1}]' + \varepsilon_t \tag{9}$$

where

$$\mathbf{A} = \begin{bmatrix} 1 & -\alpha_{12} \\ 0 & 1 \end{bmatrix}, \mathbf{B}(L) = \begin{bmatrix} -(1-\pi) & \alpha_{12}(1-\pi) & 0 & 0 \\ 0 & 0 & \beta_{21}(L) & \beta_{22}(L) \end{bmatrix}$$

By inverting **A** we can obtain the following reduced form for y_t

$$Dy_t = -(1-\pi)[y_{t-1} - \alpha_{12} x_{t-1}] + \alpha_{12}\beta_{21}(L) Dx_{t-1} + \alpha_{12}\beta_{22}(L) Dy_{t-1} \tag{10}$$

which has the following crucial features:

— it includes variables dimensionally coherent and stationary;

— therefore usual asymptotic properties apply to the estimators of its parameters;

— the long-run equilibrium relation is modelled by the error correction mechanism and the short-run dynamics are modelled by the distributed lags of the differences of the variables included in the cointegrating vector in a way consistent with the long-run behaviour.

In terms of econometric practice this procedure implies first the estimation of α_{12} and then the empirical specification of eq. (10). The latter step does not constitute a problem since the «general to specific» strategy can be applied in a straightforward way. Therefore we are left only with the problem of the estimation of α_{12}.

Stock (1987) has shown that, under the hypothesis of cointegration, the cointegrating vector can be «superconsistently» estimated by OLS: in our example $\hat{\alpha}_{12}-\alpha_{12}$ is $O_p(T^{-1})$ instead of the usual $O_p(T^{1/2})$.

To illustrate this result, briefly consider equation (7). The error term u_{1t} contains all the omitted dynamics, which can be reparameterized in terms of the variables $I(0)$ included in (10); therefore, α_{12} is consistently estimated despite the complete omission of dynamics.

We have

$$\hat{\alpha}_{12}-\alpha_{12}=(\Sigma x_t^2)^{-1}\,\Sigma x_t u_{1t} \tag{11}$$

and given that, under the hypothesis of cointegration, u_{1t} is $I(0)$ while x_t is $I(1)$ it follows that $(1/T)\,\Sigma x_t^2$ is $O_p(T)$ while $(1/T)\,\Sigma x_t u_{1t}$ is $O_p(1)$, therefore $\hat{\alpha}_{12}-\alpha_{12}$ is $O_p(T^{-1})$ and there exists a finite T^* such that for each $T>T^*$ the difference between the estimated value of the parameter and its true value is a quantity of an order smaller than $1/T$.

In theory, then, the problem of the estimation of the cointegrating vector is solved in the most straightforward vay.

However several problems affect this procedure:

— Monte-Carlo simulations (Banerjee, Dolado, Hendry and Smith (1986)) revealed that the value of T^* is definitely larger than the usual sample size available for empirical work. Moreover, if seasonal data are used, two additional problems need to be considered. The first concerns the detection of the order of integration of variables: in fact the choice of the appropriate order of differentiation of variables is not determined. The second is related to the power of the statistics for cointegration, which is affected by the non-uniqueness of unit roots;

— the cointegrating regression delivers one equilibrium relation which, when the variables involved are more than two, may not be unique. This problem could be overcome by using the Johansen (1988) procedure which delivers a statistical test for the number of cointegrating vectors and numerical values for the parameters involved. However, it is usually very difficult to give an economic interpretation to the outcome of this procedure in the sense that it is not clear which criterion should be used to select the cointegrating relation relevant to the economic problem at hand;

— structural breaks in the behaviour of the series within the sample might render the task of discovering the order of integration of a variable very difficult. In general, a stationary variable with some structural breaks could be identified as non-stationary testing procedures (Rappoport and Reichlin (1989), for example, have recently investigated the difficulties in distinguishing between a segmented deterministic trend and a stochastic trend).

Judged against this series of problems, the use of the «general to specific» procedure has a number of advantages, namely:

— the appropriate differencing is suggested by the data;

— the property of balance of the equation can be checked ex-post by looking at the level of integration of the variables included in the final specification. In particular, the fact that the ECM term is stationary would take care of above mentioned objection by Kelly;

— the cointegrating vector is uniquely determined by the data;

— Monte-Carlo experimentation has proved that the long-run coefficients obtained by means of this procedure are less affected by small sample bias than the coefficients obtained from the first step of the Granger-Engle procedure (Banerjee *et al.* (1986)).

On the basis of the foregoing considerations, the following empirical analysis is designed to illustrate both the problems encountered in applying the different methodologies of dynamic specification to the Italian money demand and more general problems potentially affecting the whole class of feedback models.

2. Simultaneous specification of long-run equilibrium and short-run dynamics in a feedback model of the Italian money demand

Following the procedure discussed in Section 1.1., we start from the estimation of the following general baseline model

$$(m-p)_t = c + \sum_{i=1}^{5} \alpha_i (m-p)_{t-i} + \sum_{i=0}^{5} \beta_i y_{t-i} + \sum_{i=0}^{5} \delta_{1i} R^b_{t-i} +$$

$$+\sum_{i=0}^{5} \delta_{2i} R^m_{t-i} + \sum_{i=0}^{9} \mu_i p_{t-i} + \Theta_1 Q_1 + \Theta_2 Q_2 + \Theta_3 Q_3 + \mu_t \tag{12}$$

where the variables are defined as follows:

m = (log of) end-of-period stock of M2 held by the public

p = (log of) GDP deflator

y = (log for) GDP

R^m = weighted average of post-tax yields of the components of M2. The weights are determined by the end of period outstanding stocks of each component

R^b = representative yield of alternative assets to M2, given by the yield of government bonds (BTP) before 1974.1 and by an average of yields of government bonds (BTP) and Treasury bills after 1974.1. The weights are determined by the end of period outstanding stocks

Q_i = seasonal dummies.

The data used are quarterly, seasonally unadjusted, from 1962.1 to 1986.2. Data sources are: Banca d'Italia, «Bollettino statistico», and ISTAT, «Supplemento al Bollettino mensile di statistica», various issues.

The underlying theoretical model of the demand for money is standard, with a scale variable and the set of relevant yields on money and alternative assets as determinants of real money holdings. Nine lags of the price level are included to allow for a potential fifth lag effect of the annual rate of inflation.

In Table 1 three successive steps of the reduction process are reported. The first (equation 1) represents an intermediate stage, where all the exclusion restrictions are imposed and tested against the baseline model in equation (12), yielding a value for the $F(18, 52)$ of 0.96. Notice that at this stage dynamics only of the first, fourth and fifth order for all variables are present and the difference restrictions on the price level are supported by the data. The implication of these latter restrictions is that the annual inflation rate (π_t) is the only variable involving prices which is relevant to the equation. Therefore, the homogeneity of degree one of nominal money to the price level and consequently the choice of real money balances as the dependent variable are supported.

The battery of diagnostic tests reported in the table shows that equation 1 is a congruent reprentation of the DGP. In particular, we have used the Jarque-Bera test for normality, two tests for heteroscedasticity (the Engle's ARCH test for sixth-order autoregressive conditional

TABLE 1 - *Simultaneous Specification of Long-Run Equilibrium and Short-Run Dynamics*

Sample period: 1964.2 to 1986.2

Equation 1: Modelling $(m-p)_t$ by OLS (st. errors in parentheses)

$$\begin{aligned}(m-p)_t = {} & \underset{(0.06)}{0.826}\,(m-p)_{t-1} + \underset{(0.06)}{0.903}\,(m-p)_{t-4} - \underset{(0.075)}{0.689}\,(m-p)_{t-5} + \\ & + \underset{(0.08)}{0.117}\,y_t - \underset{(0.09)}{0.114}\,y_{t-1} - \underset{(0.08)}{0.144}\,y_{t-4} + \underset{(0.08)}{0.101}\,y_{t-5} + \\ & - \underset{(0.001)}{0.006}\,R^b_{t-1} + \underset{(0.0015)}{0.003}\,R^b_{t-4} + \underset{(0.002)}{0.005}\,R^b_{t-5} + \underset{(0.003)}{0.007}\,R^m_{t-1} + \\ & - \underset{(0.003)}{0.012}\,R^m_{t-4} - \underset{(0.001)}{0.009}\,\pi_t + \underset{(0.001)}{0.008}\,\pi_{t-1} - \underset{(0.001)}{0.001}\,\pi_{t-4} + \\ & + \underset{(0.001)}{0.001}\,\pi_{t-5} - \underset{(0.01)}{0.012}\,Q_1 - \underset{(0.005)}{0.013}\,Q_2 - \underset{(0.005)}{0.01}\,Q_3\end{aligned}$$

$R^2 = 0.99$ $\sigma = 0.0103$ DW = 2.16 RSS = 0.007427
Normality Chi^2 (2) = 4.35 AR 1 — 6 $F[6,63] = 1.65$
ARCH 6 $F[6,57] = 0.14$ $X_i^2\ F[35,33] = 1.00$
$X_j\ F[21,47] = 1.72$ RESET $F[3,67] = 0.30$

Equation 2: Modelling $D(m-p)_t$ by OLS

$$\begin{aligned}D(m-p)_t = {} & \underset{(0.058)}{0.936}\,D(m-p)_{t-4} - \underset{(0.072)}{0.213}\,D_4(m-p)_{t-1} + \underset{(0.07)}{0.149}\,D_4 y_t - \underset{(0.07)}{0.120}\,D_4 y_{t-1} + \\ & - \underset{(0.001)}{0.003}\,D_3 R^b_{t-1} - \underset{(0.002)}{0.005}\,D_4 R^b_{t-1} + \underset{(0.003)}{0.012}\,D_3 R^m_{t-1} - \underset{(0.001)}{0.009}\,D\pi_t + \\ & + \underset{(0.02)}{0.035}\,(m-p)_{t-1} - \underset{(0.027)}{0.035}\,y_{t-1} + \underset{(0.002)}{0.002}\,R^b_{t-1} - \underset{(0.002)}{0.005}\,R^m_{t-1} + \\ & - \underset{(0.001)}{0.001}\,\pi_{t-1} - \underset{(0.004)}{0.007}\,Q_1 - \underset{(0.005)}{0.011}\,Q_2 - \underset{(0.005)}{0.007}\,Q_3\end{aligned}$$

$R^2 = 0.95$ $\sigma = 0.0102$ DW = 2.16 RSS = 0.007635

Long-run solutions[1]:

$$(m-p) = \underset{(0.013)}{1.00}\,y + \underset{(0.09)}{0.143}\,R^m - \underset{(0.06)}{0.057}\,R^b + \underset{(0.013)}{0.029}\,\pi$$

$$\text{ECMGS} = (m-p) - \{1.00\,y + 0.143\,R^m - 0.057\,R^b + 0.029\,\pi\}$$

[1] The standard errors of the long-run solutions are calculated with the formula reported in A. BANERJEE, J. GALBRAITH and J. DOLADO (1988), p. 7.

(follows)

Table 1 - *(continued)*

Equation 3: Modelling $D(m-p)_t$ by OLS

$$
\begin{aligned}
D(m-p)_t = {} & \underset{(0.04)}{0.918}\, D(m-p)_{t-4} - \underset{(0.05)}{0.194}\, D_4(m-p)_{t-1} + \underset{(0.067)}{0.142}\, D_4 y_t - \underset{(0.064)}{0.115}\, D_4 y_{t-1} \\
& - \underset{(0.001)}{0.003}\, D_3 R^b_{t-1} - \underset{(0.0016)}{0.005}\, D_4 R^{b-1}_t + \underset{(0.002)}{0.012}\, D_3 R^m_{t-1} - \underset{(0.001)}{0.009}\, D\pi_t + \\
& + \underset{(0.0067)}{0.035}\, \mathrm{ECMGS}_{t-1} - \underset{(0.003)}{0.009}\, Q_1 - \underset{(0.003)}{0.013}\, Q_2 - \underset{(0.003)}{0.009}\, Q_3
\end{aligned}
$$

$R^2 = 0.948$ $\sigma = 0.00999$ DW = 2.18 RSS = 0.00769
Normality Chi^2 (2) = 6.25 AR 1 — 5 $F[5,71] = 1.25$
ARCH 4 $F[4,68] = 0.13$ X_i^2 $F[21,54] = 2.70$
RESET $F[1,75] = 4.68$
Structural stability: see figures 1-2

heteroscedasticity and the White's test for heteroscedasticity, X_i^2), the Lagrange Multiplier test for autocorrelation of the residuals up to the sixth order (AR 6), and two tests for functional misspecification ($X_i X_j$ for misspecification due to the product of the regressors, and the Ramsey RESET test for the correct specification of the linear functional form against quadratic and cubic alternatives)[4].

The unrestricted ECM form and the final restricted ECM form are reported as equations 2 and 3 respectively.

Equation 2 has been obtained from 1 by imposing the difference restrictions suggested by the data; the $F(3,70)$ test for these restrictions is 0.66. The only difference between equations 2 and 3 consists in the imposition of the ECM term, which does not cause any change in the values of the estimated coefficients by virtue of the principle of conditioning the likelihood function.

Several comments on the final form are in order. First, the dependent variable has been respecified in first differences but very similar results both in terms of the long-run solutions and overall performance of the equation are obtained when fourth-order differences of real money balances are used as the dependent variable. Second, the resulting short-run dynamics are complicated, involving several orders of differencing, but all the signs are in line with theoretical a priori beliefs. Third, the long-run solution is characterized by unit elasticity of real money holdings to income, different semi-elasticities to the alternative interest rate R^b and the own rate of return R^m, and a positive elasticity to inflation.

The different semi-elasticities of $m-p$ to R^b and R^m reveal an asymmetric behaviour of agents with respect to movements in the two relevant

[4] For a detailed description of these tests, see D.F. Hendry (1989).

interest rates; the restriction that the differential $(R^b - R^m)$ was the only relevant variable in the regression has been rejected throughout the reduction procedure, resulting in a substantial increase in the standard error of the regression. In particular, the magnitude of the semi-elasticity with respect to R^m is more than twice the semi-elasticity with respect to R^b. This result, and the positive sign of the static long-run elasticity to the inflation rate, are two features of the final specification which seem difficult to justify on theoretical grounds.

A possible interpretation could be provided by looking at our final specification as the reduced form of a truly feedforward model. Consider the following very simplified representation of a feedback model

$$(m-p)_t = b_1 y_t - b_2 R^m_{t-1} + b_3 R^m_{t-1} + b_4 \pi_{t-1} + \epsilon_t \tag{13}$$

in which the pattern of the coefficients is consistent with our empirical findings

$$0 < b_2 < b_3 \text{ and } b_4 > 0$$

Suppose that the underlying model is a feedforward one, specified as follows

$$(m-p)_t = b_1 y_t - \alpha (R^b - R^m)^e_t + u_t \tag{14}$$

where the relevant interest rate variable is the differential $R^b - R^m$ at time t expected as of $t-1$, and there is no separate inflation effect.

If the expectations generating process for R^b and R^m can be represented as

$$R^{be}_t = \beta_1 R^b_{t-1} + \beta_2 \pi_{t-1} \qquad \text{with } \beta_i > 0 \tag{15}$$

$$R^{me}_t = \delta_1 R^m_{t-1} + \delta_2 \pi_{t-1} \qquad \text{with } \delta_i > 0 \tag{16}$$

then (13) can be interpreted as the reduced form of (14)-(16) and the following restrictions would hold

$$b_2 = \alpha\beta_1, \quad b_3 = \alpha\delta_1, \quad b_4 = \alpha(\delta_2 - \beta_2) \tag{17}$$

Therefore, our empirical results could be generated from the above model if $\delta_1 > \beta_1$ and $\delta_2 > \beta_2$.

The overall performance of the equation is good, according to the reported diagnostic tests, and the standard error of the regression is around 1 per cent.

However, the recursive stability analysis[5] reveals three major breaks

[5] See Figures 1 and 2 for one-step recursive residuals and one-step recursive Chow tests.

occurred within the sample at the beginning of 1970, in 1975 and at the end of 1983. Again, the explicit consideration of expectational variables could improve the performance of the equation under this respect, since the 1970 break occurred immediately after monetary authorities abandoned a four-year perfect interest rate stabilization policy and the 1975 break coincides with a dramatic change in the stochastic process generating inflation.

TABLE 2 - *Tests on the Order of Integration of Variables*

VARIABLE	$D(m-p)$	$D\pi$	$D_4 y$	$D_4 R^b$	$D_3 R^b$	$D_3 R^m$	ECM
DF	—1.08 (—10.2)	—.91 (—8.67)	—.19 (—3.07)	—.20 (—3.13)	—.24 (—3.51)	—.19 (—3.06)	—.059 (—1.59)
ADF	—.32 (4) (—2.28)	—	—.25 (4) (—3.13)	—.24 (5) (—3.20)	—.40 (1) (—6.61)	—.23 (4) (—3.20)	—.02 (2) (—1.09)

Notes

[1] DF is the Dickey-Fuller test, its 1% critical value here is —4.07, its 5% critical value is —3.37 (R.F. ENGLE, C.W.J. GRANGER (1987), p. 269, table II).

[2] In each cell the coefficient on the relevant lag level is reported together with its *t*-statistic.

[3] ADF is the augmented Dickey-Fuller its 1% critical value is 3.77, its 5% critical value is 3.17 (R.F. ENGLE, C.W.C. GRANGER (1987), p. 269, table II). In each cell the coefficient on the relevant lag level is reported together with its *t*-statistic and the number of lags in the dependent variable necessary to obtain white noise residuals.

Finally, as far as the balance of the equation is concerned, Table 2 shows that all the variables in the final specification are close to stationarity, with the only notable exception of the ECM term (Figure 3). This finding is to be related to our previous discussion of Kelly's (1985) observation that the long-run elasticities derived from feedback models may not capture deep behavioural parameters, because of the presence of expectational parameters. A non-stationary ECM term tends to support Kelly's view, since in this case the counterargument put forward by Hendry and Neale (1988) does not apply.

3. TWO-STEP SPECIFICATION OF LONG-RUN EQUILIBRIUM AND SHORT-RUN DYNAMICS OF THE DEMAND FOR MONEY

In this section the Engle-Granger two-step procedure outlined in Section 2 is applied and results are reported in Table 3.

The estimate of the cointegrating regression between the same variables studied in the preceding analysis delivers a CRDW of 0.943 and rejects the hypothesis of non-stationarity of the residuals from this regression, shown in Figure 4 as ECMGE.

TABLE 3 - *Two-Step Specification of Long-Run Equilibrium and Short-Run Dynamics*

Sample period: 1964:2 to 1986:2 (st. errors in parentheses)

A) *First step*

$$\underset{}{m-p} = \underset{(0.01)}{1.223\, y} - \underset{(0.009)}{0.002\, R^m} - \underset{(0.005)}{0.017\, R^b} + \underset{(0.003)}{0.018\, \pi}$$

$R^2 = 0.99 \qquad \sigma = 0.0947 \qquad DW = 0.943$

B) *Second step*

$$\begin{aligned} D(m-p)_t = {} & \underset{(0.05)}{0.915\, D(m-p)_{t-4}} + \underset{(0.035)}{0.050\, D_4(m-p)_{t-1}} + \underset{(0.077)}{0.143\, D_4 y_t} - \underset{(0.076)}{0.149\, D_4 y_{t-1}} + \\ & - \underset{(0.0015)}{0.004\, D_3 R^b_{t-1}} - \underset{(0.0017)}{0.002\, D_4 R^b_{t-1}} + \underset{(0.0026)}{0.010\, D_3 R^m_{t-1}} - \underset{(0.0009)}{0.010\, D\pi_t} + \\ & - \underset{(0.019)}{0.030\, \mathrm{ECMGE}_{t-1}} - \underset{(0.003)}{0.005\, Q_1} - \underset{(0.004)}{0.002\, Q_2} - \underset{(0.003)}{0.003\, Q_3} \end{aligned}$$

$R^2 = 0.93 \qquad \sigma = 0.0115$	DW = 2.037 RSS = 0.01004
Normality Chi^2 (2) = 16.98	AR 1 — 5 $F[5,70] = 0.67$
ARCH 4 $F[4,67] = 0.13$	$X_i^2\, F[21,53] = 2.21$
RESET $F[1,74] = 8.87$	

The estimated long-run equilibrium coefficients present the following features: *a*) an elasticity of 1.22 to income, slightly greater but consistent with the unitary value found with the «general to specific» approach; *b*) a much smaller semi-elasticity to the alternative interest rate R^b (—0.017) and an almost zero coefficient on R^m; *c*) a positive but smaller long-run responsiveness to the inflation rate (0.018).

However, recursive estimation (Figures 5-8) shows that such estimates, with the only possible exception of the income elasticity, suffer from a high degree of instability over the sample, making the results from this first stage of the procedure not entirely reliable. In particular, the coefficient on R^m is positive for a substantial part of the sample before becoming insignificant and the coefficient on the inflation rate becomes significantly positive only towards the end of the sample.

In the second step, the ECMGE term is included in the baseline equation (1) and the usual process of reduction is implemented to identify the short-run dynamics. The final outcome is a specification whose performance is very close to the «general to specific» equation.

Although the standard error of the regression is slightly higher, the performance in terms of diagnostics is very similar and the recursive stability analysis (Figures 9 and 10) shows the same pattern of structural breaks. Finally, the coefficient on the ECMGE term, although negative, is not significant and too small to favour a feedback error-correcting interpretation.

A general criticism to the two-step procedure can be related to the fact that the maximum number of cointegrating vectors between our five variables is four and it is not clear whether the first step of the Granger-Engle procedure selects one of them or a linear combination of all the existing cointegrating vectors. To tackle this problem, we adopt the procedure recently proposed by Johansen (1988), briefly described in the *Appendix*. This procedure delivers a statistical test for the number of existing cointegrating vectors between N variables and estimates for their coefficients. The test has a non-standard distribution and appropriate critical values are reported in Johansen and Juselius (1989). The results from this procedure are reported in Table 4.

TABLE 4 - *The Johansen Procedure*

Sample period: 1964.2 to 1986.2

Number of Coint. Vec.	Stat. Test (crit. val.)	Coefficients in the Associated Cointegrating Vector				
		y	π	R^m	R^b	$m-p$
1	106.98 (78.9)	−1.56	−0.24	2.72	−1.64	−1.00
2	66.23 (56.4)	3.98	0.164	−1.37	1.61	−1.00
3	35.66 (37.6)	2.23	0.025	−0.02	−0.042	−1.00
4	16.42 (22.2)	1.33	0.03	0.013	−0.022	−1.00
5	6.66 (10.7)					

The statistical test shows that the null of at most two cointegrating vectors cannot be rejected at the 97.5 per cent cut-off point.

The first three cointegrating vectors contain values of the long-run equilibrium coefficients either of the wrong sign or of implausible magnitudes, whereas the fourth one presents elasticities to income and inflation sufficiently close to those derived from the Granger-Engle and the «general to specific» procedures. However, the semi-elasticities to interest rates, although of the correct sign, are substantially different. The conclusion from this procedure is that, although it may provide a useful check of the univariate procedure, it adds very little economic insight.

4. A NOTE ON A RECENTLY PROPOSED DYNAMIC SPECIFICATION STRATEGY FOR ITALIAN MONEY DEMAND

Papi (1988) has recently carried out a detail analysis of the feedback specification of the Italian money demand in a «general to specific»

TABLE 5

Sample period: 1964.2 to 1986.2

Modelling $D(m-p)_t$ by OLS

$$
\begin{aligned}
D(m-p)_t = {} & \underset{(0.052)}{0.946}\,(m-p)_{t-4} - \underset{(0.069)}{0.721}\,(m-p)_{t-5} + \underset{(0.078)}{0.138}\,y_t - \underset{(0.101)}{0.326}\,y_{t-1} \\
& - \underset{(0.077)}{0.159}\,y_{t-4} + \underset{(0.071)}{0.122}\,y_{t-5} + \underset{(0.001)}{0.002}\,R^b_{t-4} - \underset{(0.001)}{0.006}\,D_4R^b_{t-1} \\
& + \underset{(0.003)}{0.007}\,R^m_{t-1} - \underset{(0.002)}{0.012}\,R^m_{t-4} - \underset{(0.001)}{0.001}\,\pi_{t-1} - \underset{(0.001)}{0.009}\,D\pi_t \\
& - \underset{(0.054)}{0.186}\,(m-p-y)_{t-1} - \underset{(0.009)}{0.009}\,Q_1 - \underset{(0.005)}{0.011}\,Q_2 - \underset{(0.005)}{0.007}\,Q_3
\end{aligned}
$$

$R^2 = 0.948$ $\sigma = 0.01023$ DW = 2.14 RSS = 0.00763
Normality Chi^2 (2) = 4.79 AR 1—5 $F[5,68] = 1.25$
ARCH 4 $F[4,65] = 5.11$ $X_i^2\ F[29,43] = 1.31$
RESET F [1,72] = 5.76

Long-run solutions: $(m-p = 1.00\,y + 0.128 \quad R^m = -0.051 \quad R^b + 0.026\,\pi$

framework. The main feature of his procedure is that a particular form for the error correction term is imposed since the beginning of the specification search. This term is the reciprocal of velocity: $(m-p-y)$, which has been successfully used by Hendry in several studies of the demand for money in the United Kingdom (Hendry and Mizon (1978), Hendry (1985)).

Several comments are in order. In the first place, the overall performance of the equation is satisfactory, although there is some sign of autoregressive heteroscedasticity and functional misspecification. The structural stability analysis (Figures 11 and 12) shows the same pattern as previous specifications.

The long-run solutions of the model are very close to those obtained from the «general to specific» procedure implemented in Section 2. However, the error correction term is not formed by the long-run equilibrium, and the specification captures only the long-run dynamics without modelling the short-run dynamics around it.

In fact, the coefficient of -0.186 does not represent the effect of the levels of $(m-p)$ on $D(m-p)$ which can be derived from the estimated equations as $+0.039$, very close to the $+0.035$ estimate obtained in Section 2. Therefore, the interpretation of the significance of the imposed error correction term as reflecting a feedback error correcting behaviour of agents is not allowed, since it would require a negative coefficient on the linear combination of variables capturing the long-run equilibrium.

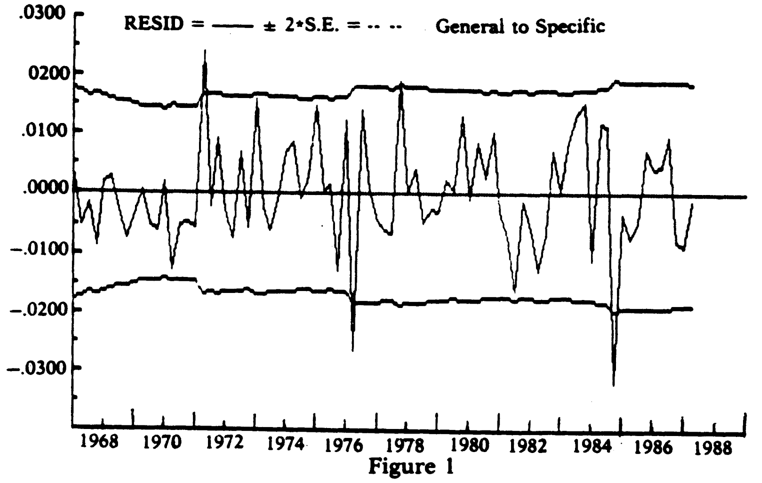

Figure 1

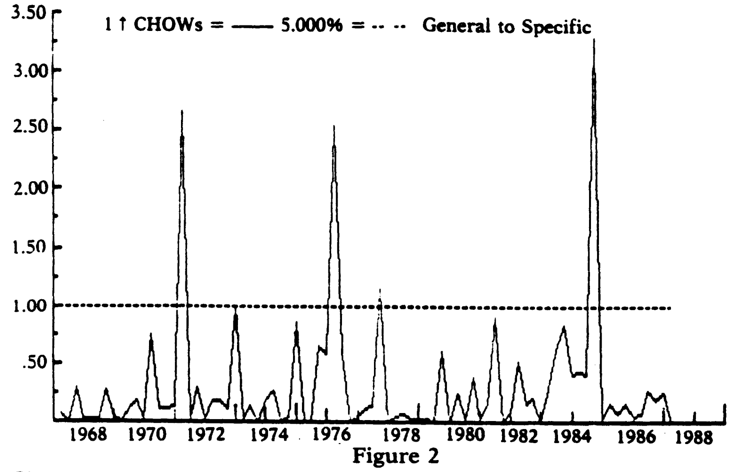

Figure 2

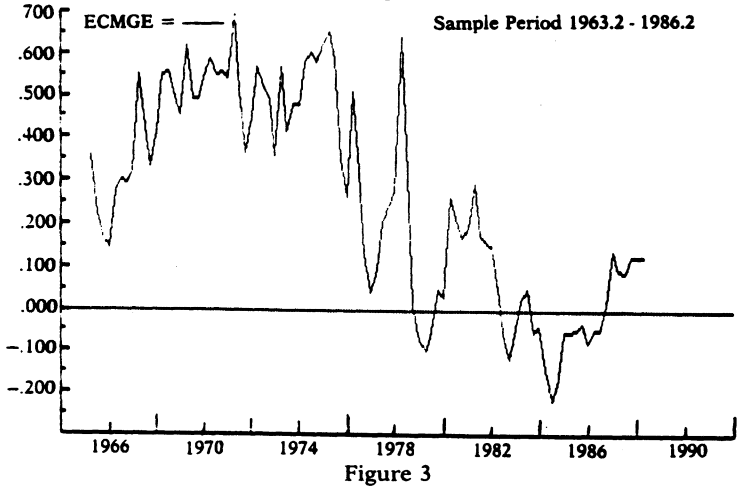

Figure 3

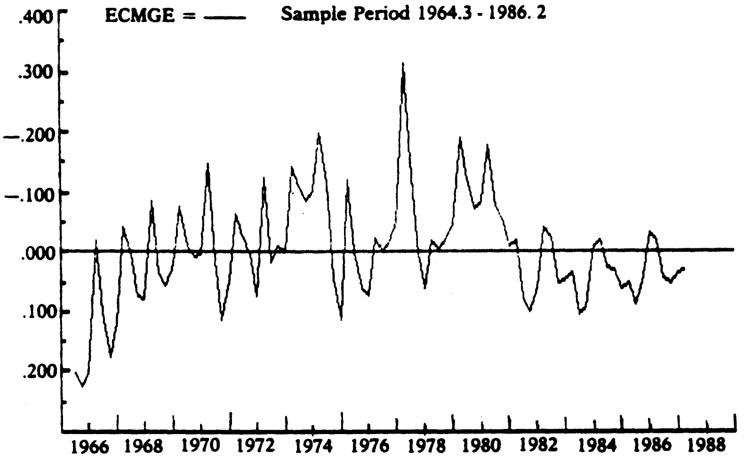

Figure 4

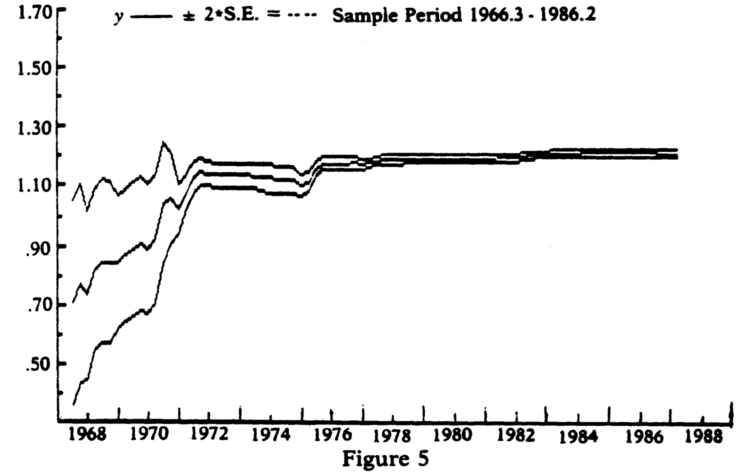

Figure 5

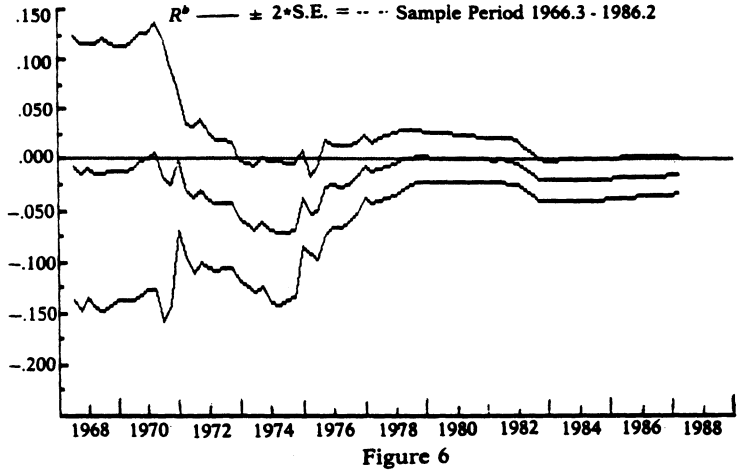

Figure 6

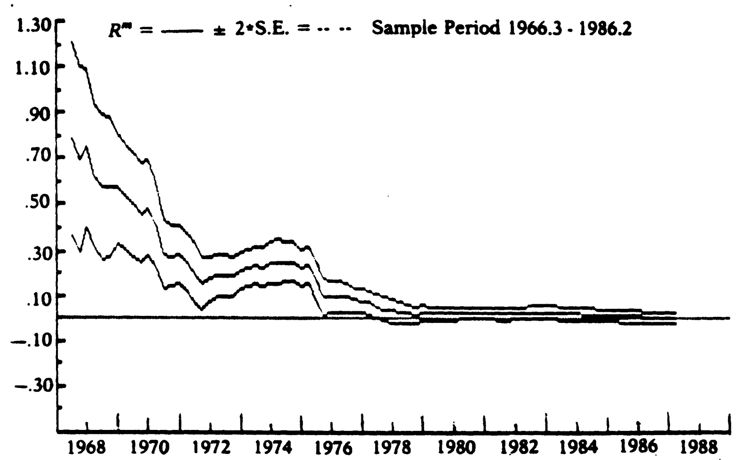

Figure 7

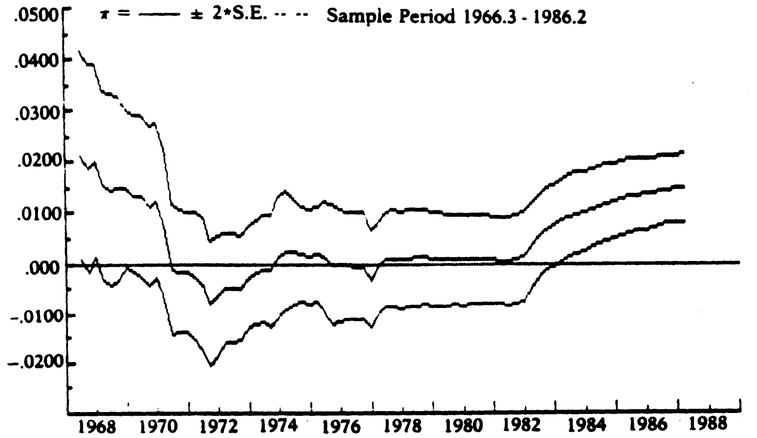

Figure 8

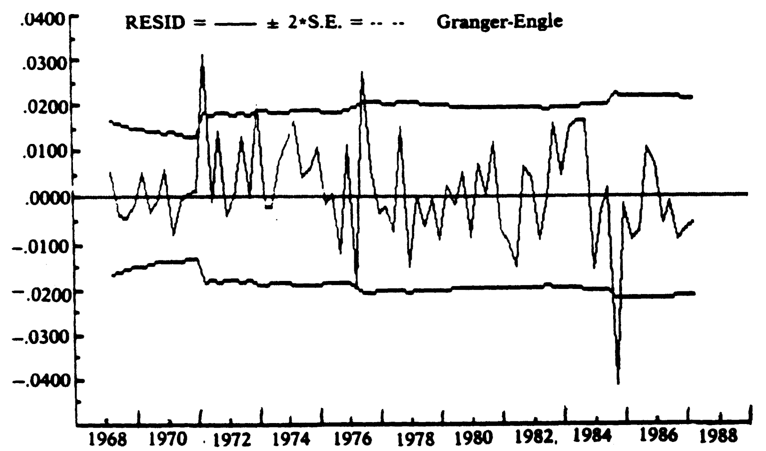

Figure 9

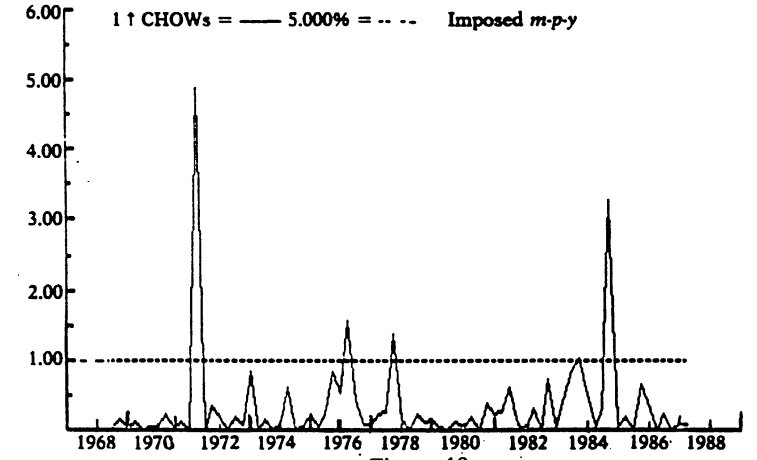

Figure 10

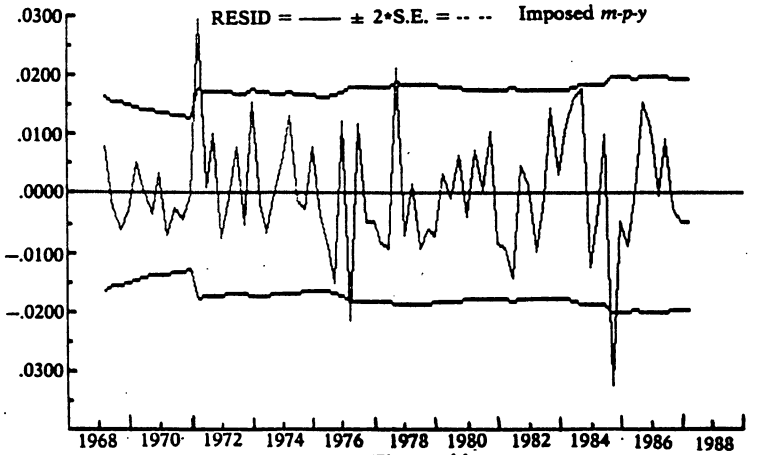

Figure 11

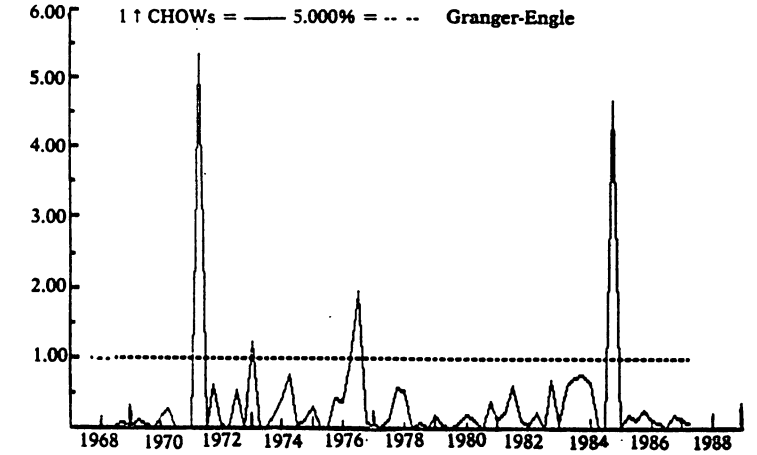

Figure 12

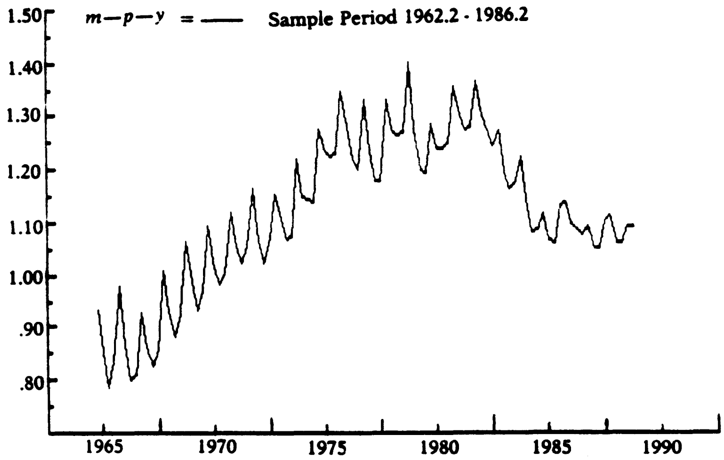

Figure 13

Moreover, the property of balance of the equation is not ensured: levels and differences of the same variables are simultaneously included, together with an *a priori* linear combination of the variables.

Kelly's argument is potentially relevant in this context, in view both of the structural instability of the equation and the clear non-stationarity of $(m-p-\dot{y})$, as Figure 13 shows.

5. Conclusions

The specification and estimation of empirical models based on the assumption of feedback behaviour has become a general practice in applied econometrics.

Several potential problems affect feedback specification. They are mainly related to the existence of forward-looking agents whose behaviour is determined by expectations and cannot be satisfactorily modelled without specifying the expectations generating process.

However, given any feedback specification, it is possible to identify it as a true structural form or a reduced form of a feedforward model by means of structural stability analysis and by checking the property of cointegration between the variables included in the model to capture the long-run solution.

When a feedback model is specified for the demand for money in Italy the use of different specification techniques yields results which are similar, but consistently affected by structural instability and non-stationarity of the long-run solution. Therefore, a potentially more successful specification strategy seems to require the explicit modelling of the process generating expectations.

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APPENDIX

THE JOHANSEN PROCEDURE

Johansen (1988) considers a VAR

$$A(L)\,X_t = \epsilon_t \tag{A.1}$$

where $A(L)$ is a k-th order polynomial, X_t is a $p \times 1$ vector of $I(1)$ variables and ϵ_t is NIID $(0, \Omega)$.

The model can be reparameterized as

$$DX_t = \cap_1 DX_{t-1} + \dots + \cap_{k-1} DX_{t-k+1} + \cap_k X_{t-k} + \epsilon_t \tag{A.2}$$

where $\cap_i = -I + A_1 + \dots + A_i \qquad i = 1$ to K.

If we call $\cap_k = A$, we have that A will be an $P \times P$ matrix and the number of distinct cointegrating vectors which exist between the variables included in X will be given by the rank of $\cap$, r.

Since X consists of variables which must be differenced once in order to become stationary, then, at most, r must be equal to $p-1$. If we define two matrices $\boldsymbol{\alpha}$, $\boldsymbol{\beta}$ both $p \times r$ and we express $\mathbf{A} = \boldsymbol{\alpha\beta}'$, we have that the rows of $\boldsymbol{\beta}$ form the r distinct cointegrating vectors and the space spanned by $\boldsymbol{\beta}$ can be estimated.

If the effects of $(DX_{t-1}, \dots DX_{t-k+1})$ are partialed out from DX_t and X_{t-k} by regression to obtain residuals R_{0t} and R_{mt}, respectively, then it is possible to compute the second moments of all these residuals, denoted S_{00}, S_{0m}, S_{mm} where

$$S_{ij} = T^{-1} \sum_{t=1}^{T} R_{it} R_{jt} \text{ for } i, j = 0, m$$

the likelihood ratio test statistic for the hypothesis that there are at most r cointegrating vectors is

$$-T \sum_{r+1}^{N} \log((1 - \mu_i) = Q$$

where μ_i are the smallest eigenroots obtained by solving

$$|\mu S_{mm} - S_{m0} S_{00}^{-1} S_{0m}| = 0$$

the corresponding eigenvectors $\beta(\epsilon_1, \ldots \epsilon_r)$ normalised by $\beta' S_{mm} \beta = I$ are the cointegrating vectors (if there are any).

The critical values for the statistic Q are tabulated in Johansen and Juselius (1989).

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