Money, policy regimes and economic fluctuations

by

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Abstract

Part I deals with the estimation of money demand functions. Several non-structural interpretations of the conventionally estimated functions are surveyed and discussed (*Chapter 1*). An application to Italian data is then presented, focusing on two such interpretations.

First (*Chapter 2*), the role of expectations in determining money demand behaviour is assessed. Since monetary policy regimes have a direct effect on the time-series properties of interest rates, the identification of clear regime changes may provide a powerful test of forward-looking models of money demand. An expectations model is constructed, which is stable in the face of the Italian monetary policy regime change in 1970, when traditional backward-looking money demand functions show remarkable instability.

Second (*Chapter 3*), the existence of multiple long-run relations among the variables relevant to money demand is shown to create problems for the interpretation of single-equation estimates. To obtain a satisfactory specification of the long-run relations and the short-run dynamics of the system around equilibrium, a sequential procedure is devised and applied.

In *Part II*, the controversy between "real" and "monetary" theories of fluctuations is examined (*Chapter 4*). A "monetary" equilibrium model of the cycle is constructed, extending the original Lucas "island" framework to allow for a powerful role for stabilization policy. The implications of alternative monetary policy regimes are derived and tested on U.S. data, comparing two periods (1922-1940 and 1952-1968) with a different policy stance.

Chapter 5 investigates the relative importance of the "money" and "credit" channels of monetary transmission for Italy in the 1982-1994 period, using a structural *VAR* methodology. Monetary policy is effective, though not through a "credit channel", and independent disturbances to credit supply have sizeable real effects.

In *Chapter 6* the focus is shifted to anticipated fiscal policy actions and their effect on consumption. A long series of pre-announced income tax changes is examined for the U.K.. Consumption reacts to such fiscally-induced disposable income changes only at the implementation dates.

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Introduction

This thesis deals with several topics in macroeconomics and applied monetary economics.

Part I is concerned with the estimation and economic interpretation of money demand equations. Despite being a long-standing topic in applied monetary economics, the empirical analysis of money demand has recently attracted renewed attention, given the poor forecasting performance of conventional econometric equations. The well-known repeated episodes of instability of estimated functions occurred during the 1970s and 1980s in many countries -and particularly in the U.S.- stimulated the empirical research in various directions. On the one hand, some authors attributed instability to the omission of relevant variables, usually related to financial innovation, and to the dynamic mis-specification of the empirical models. For example, Baba, Hendry and Starr (1992) argued that the inclusion of measures of risk and return on long-term bonds and a more careful specification of the equation's dynamic structure may eliminate the main episodes of instability in empirical money demand functions for the U.S., widely known as the "missing money" puzzle (1974-76), the "great velocity decline" (1982-83) and the "M1 explosion" (1985-86).

On the other hand, a radical criticism of the capability of single-equation models of yielding valuable information on agents' behavioural characteristics has been put forward by Cooley and LeRoy (1981). The recognition of the simple fact that the money demand function naturally belongs to a system of (maybe complex) relationships among monetary aggregates, income, inflation and interest rates, implies that the parameters delivered by conventional money demand equations may not be true behavioural parameters, depending in various ways onto the processes generating money demand determinants. As a consequence, instability problems may be unrelated to shifts in the underlying money demand parameters, being caused instead by changes in the time-series behaviour of other variables.

This general consideration provides the unifying theme of the three chapters of Part I, the first of which is devoted to the discussion of various *non-structural* interpretations of conventional money demand regressions. In particular, two such interpretations directly motivate the empirical investigations of the following chapters.

Firstly, 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 expectational 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 money demand elasticities. Of course, this argument is an application of the general Lucas (1976) critique of conventional econometric models (Cuthbertson and Taylor (1990)).

Secondly, if there exist multiple long-run relations linking money balances and other relevant variables (e.g. income, interest rates, inflation), the estimates of long-run money demand parameters based on single-equation models may well be combinations of such multiple relations among the series under study. The recently developed cointegration theory (Johansen and Juselius (1990)) provides tools for estimating long-run (equilibrium) relationships in a system context, allowing also for testing of specific structural hypotheses on the economic nature of the detected relations, and may be usefully applied to tackle this problem.

We investigate the potential relevance of expectations in explaining instability of the Italian demand for M2 over the 1964-1986 period in chapter 2. Particular attention is devoted to the response of money demand to the clear change in the monetary policy regime occurred in 1970, which dramatically altered the time-series behaviour of interest rates (a point overlooked by the existing empirical literature). The correspondence between structural breaks of feedback models of money demand and sharp alterations in the processes generating (some of) the regressors is formally established by means of the superexogeneity and invariance tests proposed by Engle and Hendry (1993). Then, an explicitly forward-looking model of money demand is estimated and its stability properties in the face of changes in the prevailing monetary policy regime assessed. The results do suggest that the neglect of expectations may explain instability at times of readily perceived monetary policy changes.

In chapter 3, Italian data over a period of overall stability even of conventional feedback money demand equations (1983-1991) are used to address the issue of the economic interpretation of multiple long-run relationships in a system including M2, income, interest rates and inflation. Contrary to previous studies, which recognized the problem without providing a solution (Muscatelli (1991)), we formulate and test some explicit structural economic hypotheses on the long-run equilibrium path of the system. Then, a simultaneous system of equations is specified with a short-run dynamics consistent with the proposed economic interpretation of the estimated long-run relations. The restrictions embodied in this final structural model are then tested against the reduced form of the system.

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Chapters 4 and 5 in *Part II* address the broader issue of the effect of monetary policy actions on real variables. In chapter 4 a quite general perspective is taken, considering the role of stabilization monetary policy in the context of the debate between "real" and "monetary" theories of cyclical fluctuations. Recent theoretical developments have emphasized predominantly real explanations for business cycles, largely determined by technological shocks transmitted to the whole economy through real propagation mechanisms. In support of this view, empirical results showing the absence of Granger-causality from monetary (more generally, nominal) to real variables and the negligible role of monetary disturbances in explaining output variability are often presented (Eichenbaum and Singleton (1986), Plosser (1991)). The evidence of co-movements of real and monetary variables is then explained on the basis of a reverse causation argument.

However, such empirical findings may be reconciled with a modified version of the well-known "monetary" model of the cycle due to Lucas (1973), where an expected inflation effect on aggregate demand provides a channel for monetary policy effectiveness. When monetary policy is deliberately (end effectively) used to stabilize output, the very pattern of empirical findings mentioned above may be obtained. Therefore, the possibility of discriminating among alternative theories of fluctuations simply on the basis of such tests seems questionable. A potentially more fruitful empirical strategy could exploit changes in the policy regime: for example, a switch from a countercyclical to a fixed money rule is associated with a larger impact of nominal (monetary) innnovations on output in our extended "monetary" framework, whereas it should be totally irrelevant under a "real" view of the cycle. This strategy is applied to the U.S., comparing the quantitative importance of monetary disturbances in explaining output variability in two periods (1922-1940 and 1952-1968) characterized by a different policy stance, with monetary policy systematically used for stabilization purposes only in the postwar years. The results do not support a purely real theory of economic fluctuations.

In chapter 5 we investigate the relative importance of two channels of transmission of monetary policy actions to the real economy. The first is the traditional "money" channel, working through interest rate movements in the aftermath of a change in banks' reserves implemented by the central bank, with real effects on the interest rate-sensitive components of spending. Conversely, the "credit view" of the monetary transmission mechanism (recently emphasized for the U.S. by Bernanke and Blinder (1992) and Kashyap, Stein and Wilcox (1993)) focuses on the asset side of the banking sector balance sheet and stresses the possibility for the effectiveness of monetary policy to be enhanced if restrictions of bank creadit may not be offset (at least for a significant fraction of borrowers) by recourse to alternative sources of finance.

The empirical evidence available for the U.S. does not provide undisputed support for the credit channel of transmission; in particular, the crucial problem of the identification of movements in the amount of credit outstanding as due to demand or supply shifts has not been solved (see the opposite interpretation of very similar empirical evidence offered by Romer and Romer (1990) and Bernanke and Blinder (1992)). Our empirical analysis uses Italian data for the 1980s and early 1990s and directly addresses identification issues in the context of the structural *VAR* methodology. The results do not show a powerful role for the credit channel of transmission of policy impulses, but highlight the importance of autonomous disturbances to bank loan supply in determining fluctuations in production, therefore favouring a broader "credit view" of the links between financial aggergates and real variables.

In the final chapter 6 the focus is shifted to fiscal policy actions and to their effects on consumption expenditure. From the perspective of the rational expectations-permanent income model of consumption (Hall (1978), Deaton (1992)) only unanticipated changes in agents' real lifetime resources should be reflected in innovations in consumption, whose current level incorporates all available information on future incomes and interest rates. Therefore pre-announced variations in disposable income, induced by fiscal policy actions, should have no effect on consumption expenditure when they are actually realized, being reflected in expenditure at the announcement date.

The existence in the U.K. of a substantial lag between the announcement of changes in income taxation and their actual implementation provides an ideal set-up for testing the main implications of the rational expectations-permanent income theory. Our empirical analysis, spanning a long time period (1960-1990), yields robust findings of a positive, and quantitatively significant, reaction of consumption expenditure to fiscally-induced increases in disposable income only at the implementation date, contrasting with the implications of the theory and casting doubts on the validity of the Ricardian Equivalence proposition.

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Part I

Chapter 1

Non-structural interpretations of money demand regressions: a survey of the problems and the relevant literature.

1. *Introduction*

The theoretical foundations and empirical specification of the money demand function have always been central issues in monetary economics. At the theoretical level, in the 1960s and 1970s, they have been the focus of the early dispute between competing approaches to macroeconomics, and in particular between the "keynesian" and "monetarist" views of the transmission mechanism. At the policy level, the existence of a stable money demand function is necessary for the proper design and implementation of monetary targeting policies, widely adopted since the mid-'70s both in Europe and in the United States. For this main reason, the repeated episodes of instability of conventional estimates of the function during the 1970s and early 1980s have been viewed as particularly serious issues and have been extensively studied in the literature. Various explanations for these phenomena have been offered by several authors, mainly based on the effects of financial innovation (Judd and Scadding (1982) and Goldfeld and Sichel (1990) provide comprehensive surveys of the empirical literature).

At a more general level, the interpretation of conventional empirical equations as *structural* money demand functions has been questioned in an important contribution by Cooley and LeRoy (1981). They emphasized the difficulties in formulating equilibrium models of financial markets which allowed for the identification of structural money demand parameters. At the end of their study, Cooley and LeRoy pessimistically declared themselves "*unpersuaded by existing attempts to estimate a money demand equation, but ... unable to supply an attractive alternative*" (1981, p.834). These ideas have been expanded in the more recent literature, and various authors have developed *non-structural* interpretations for conventional equations, usually presented as aggregate money demand functions. According to these views, the money demand equation naturally belongs to a broader system of relationships between the variables under study (money, prices, income, interest rates), and the parameters of the conventionally estimated equations may not be structural parameters, being related in various ways to the processes generating the determinants of money demand. Consequently, the estimated parameters can vary if these processes alter through time, with no change in agents' behaviour or in the degree of financial evolution of the economy. In particular, the neglect of expectations in traditional empirical specifications has been considered as one potential source of instability.

The empirical analysis of chapter 2 is designed to assess the role of expectations by comparing a backward-looking and a forward-looking model of Italian money demand and evaluating their stability. The emphasis will be put on the comparative stability performance of the alternative models in the face of substantial changes in the prevailing (monetary) policy regime. In chapter 3 money demand is studied in a multivariate framework, with the focus on the identification and interpretation of the (multiple) long-run relationships among the variables analyzed.

The present chapter provides the theoretical and empirical background relevant to the empirical investigations of the next two chapters. To this aim, section 2 sketches some recent developments in the theoretical literature and recalls the main episodes of instability in the 1970s and 1980s in the U.S. and in the U.K., motivating a large part of the empirical work in this area. Section 3 presents various views on the non-structural nature of conventional money demand regressions. In particular, the role of expectation formation is discussed and the implications in terms of stability of feedback and forward-looking models of money demand are derived in section 3.3, setting the scene for the application of chapter 2. In section 3.4 the need for a multivariate approach to the study of money demand in the presence of multiple long-run relations among the variables involved in the analysis is illustrated, providing the motivation for the empirical study of chapter 3.

2. *Theoretical developments and empirical issues in the analysis of money demand.*

Theoretical developments in the study of the money demand function over the last fifteen years have come mainly from the effort of explaining some unsatisfactory features of the empirically estimated equations, namely temporal instability and implausibly long adjustment lags to movements in the determining variables. At the level of pure theory, some progress has been made along the lines set out by the classic contributions of Baumol (1952), Tobin (1956) -refined and extended by Miller and Orr (1966, 1968) and Orr (1971)- and Whalen (1966) on the transactions and precautionary motives, and of Tobin (1958) on the speculative demand for money. As Fischer (1988, p.295) notes, "[t]*heoretical work on the demand for money was a declining industry in 1975, and there has been only a brief subsequent revival*", mainly due to the work of Akerlof, Milbourne and D. Romer.[1]

In a series of papers, Akerlof (1979, 1982) and Akerlof and Milbourne (1980a,b), using a model in which agents adopt constant target-threshold monitoring rules, showed that the resulting income elasticity may be remarkably low, with velocity absorbing changes in income in the short-run, in accord with several influential empirical studies, including Goldfeld (1973). Interesting implications for the interest rate elasticity of money demand are derived by D. Romer (1986) from a continuous time overlapping generations model in which money and interest-bearing bonds coexist. In this model, which represents a development of the original Baumol-Tobin insights in a general equilibrium setting, money is necessary to purchase goods (Clower constraint condition). The consumption-saving choice of individuals is then studied together with their decision on the number, timing and size of conversions from bonds into money, given the presence of fixed transaction costs. It is shown that, in addition to the traditional channel through which changes in the interest rate affect money holdings (altering the frequency of conversions, as in the original Baumol-Tobin model), two other effects are at work. Interest rate increases positively affect money holdings through their influence on the pattern of spending between conversions and through a wealth effect, therefore tending to offset the negative effect of the increased frequency of conversions. In general, when all three channels are present, very small negative (or even

[1] Early theoretical treatments of the demand for money are reviewed by McCallum and Goodfriend (1987), Goodhart (1989b) and Goldfeld and Sichel (1990). McDonald and Milbourne (1990) survey broader developments in monetary theory. More recently, a different line of theoretical research has been pursued by Kiyotaki and Wright (1991, 1993), who try and explain, in a search-theoretic framework, why agents willingly hold non-interest bearing fiat money.

positive) elasticities of aggregate money demand to nominal interest rates can be generated.[2]

More directly useful to explain the problems of estimated money demand functions, are the recent developments in the *buffer-stock* approach to the monetary sector originated by Laidler (1982, 1984) and Goodhart (1984). Such an approach offers some explanation for at least two of the problems of empirical money demand functions based on conventional partial adjustment mechanisms: *a*) the overshooting of the interest rate, necessary to clear the money market after a change in money supply, implied by estimated short-run interest rate elasticities smaller than the long-run responses, and *b*) the implausibly long adjustment lags caused by the presence of the lagged dependent variable with an usuallý very high (though less than one) estimated coefficient. The buffer-stock view emphasizes the role of money assets as a means of payment, used in transactions on all markets. Because of this special role of money, shocks in all (goods, financial, factor) markets in which agents operate are likely to have an immediate effect on monetary flows and, given the relatively low costs of adjusting money balances with respect to other, less liquid, financial assets, agents will be willing to allow their money holdings to vary in the face of unforeseen shocks, at least in the short-run. Only subsequently will individuals reconsider their plans concerning production processes, purchases of durable goods, price setting or large portfolio reallocations. Money balances will then perform the function of a financial *buffer*, absorbing a large portion of the unexpected changes in receipts and expenditures.

According to its proponents, particularly Laidler (1984), the view of money as a buffer-stock asset has important implications for the monetary transmission mechanism. In fact, following an unexpected exogenous increase in nominal money, a state of disequilibrium would occur, determining a stream of expenditure (a *real balance effect*) which will cause movements in interest rates, output and prices, and gradually eliminate the discrepancy between money demand and supply.[3] No overshooting of the interest rate is required to reestablish equilibrium, since agents will allow money supply shocks to be absorbed initially by movements in buffer-stock holdings. Furthermore, buffer-stock models allow one to interpret conventional Goldfeld-type money demand equations -displaying

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[2] In a companion paper, D. Romer (1987) studies the effects of interest rate shocks in his model and relates the results to those obtained by Grossman and Weiss (1983) in a similar framework but with the assumption of a fixed frequency of conversions of bonds into money.

[3] This account of the transmission mechanism mainly reflects Laidler's own interpretation of the buffer stock theory. Bain and McGregor (1985) provide a comparative discussion of the buffer stock approach and other theories of the transmission mechanism. Milbourne (1987, 1988) criticizes the simple transposition at the aggregate level of the buffer stock notion of money developed at the individual level.

implausibly long adjustment lags- not as structural money demand functions but as semi-reduced forms for one of the arguments of the function.[4] Laidler (1982) favours the reinterpretation of money demand equations as price level equations, therefore explaining long adjustment lags as due to sluggishness in price level movements rather than in the portfolio adjustment process, whereas Artis and Lewis (1976) view the interest rate as the dependent variable in the equation, then emphasizing price sluggishness in financial markets rather than in goods markets. Laidler's hypothesis has been tested on U.S. data by MacKinnon and Milbourne (1988), who found no support for the view that conventional money demand equations are really semi-reduced form price equations, and Fischer and Nicoletti (1993) provided further evidence against this view on the basis of weak exogeneity tests in a cointegration framework. Several attempts at testing the validity of the buffer-stock hypothesis on the working of the money market have been performed using systems of equations. Miller (1990) finds support for the buffer-stock view from a system error-correction model applied to U.S. data for M2 in the 1959-1987 period. Lastrapes and Selgin (1994) explicitly identify money demand and supply disturbances in a bivariate vector autoregression system including both real and nominal money balances for the U.S. (1957-1991): the finding of a sizeable role of nominal M1 shocks (identified as structural money supply disturbances) in determining the short-run dynamics of real money balances is interpreted as supporting the buffer-stock view of monetary adjustment.

As already noted, the most serious problem with empirical money demand functions in the last two decades has been the poor forecasting performance displayed by conventionally specified equations both in the United States and in the United Kingdom in the mid-'70s and again in the early 1980s. In the U.S., the well-known "case of the missing money" analyzed by Goldfeld (1976) was the first episode of instability of money demand equations for the M1 aggregate. The unpredicted rise in velocity occurred in 1974/75 has been given various explanations in the literature, thoroughly surveyed by Judd and Scadding (1982). In particular, the use of incorrect definitions for the relevant determinants of money demand (especially interest rates) and financial innovations leading to a decline in the amount of money needed for transactions and investment purposes by households and firms, have been investigated as potential reasons for the instability problem. More recently, however, Rose (1985) and Baba, Hendry and Starr (1992) have shown that a more careful dynamic specification, together with the inclusion of a measure of risk to long-term bond holding (proxied by the standard deviation of the monthly holding period yield on long-term bonds)

[4] As noted by Milbourne (1988), the assumption of money supply exogeneity is essential to this argument.

in the estimated equation, can lead to substantial gains in stability over the 1974/76 period. In the U.K., in 1972/73 an increase in bank lending to the private sector, funded by a surge in bank deposits due to banks' liability management practices, determined a large underprediction of broad monetary aggregates (£M3) forecasts based on previously estimated demand for money functions. As Goodhart (1989a, p.314) concludes, "[t]*he consensus remains that the demand for money function for £M3 broke down in 1972/3, and has remained unstable ever since*".

Again in the 1980s, and not only in the U.S. and the U.K., short-run estimated equations displayed signs of instability, this time in the face of a generalized and prolonged decline in velocity. In the U.S., according to the explanation of Baba, Hendry and Starr (1992), greater variability in asset prices brought about by the monetary authorities' shift from interest rates stabilization to monetary base targeting, determined an increase in the demand for M1 for precautionary and speculative purposes. Moreover, it could be the case that previous studies of money demand, conducted in periods when the interest rates on some monetary aggregates were fixed or had upper ceilings, underestimated the interest elasticity of money balances. In the 1980s, with declining nominal interest rates and increasing financial innovation, the interest elasticity may well have increased, contributing to the poor forecasting performance of previously estimated equations (Poole (1988)).[5] In the U.K., the main reason for the decline in velocity can be found in the increase in competitiveness within the banking sector, leading to more attractive interest rates on deposits and reduced costs of borrowing for the personal sector, with a consequent surge in private sector claims upon, and indebtedness to, the banking system (see Goodhart (1989a)).

These repeated episodes of instability have badly damaged the reliability of *short-run* estimated equations, constructed in order to capture the dynamics of monetary aggregates, as a useful tool for monetary policy analysis. Indeed, several authors have argued that the estimated equations do not represent structural money demand functions, but either they may

[5] The empirical model of Baba *et al.* (1992) (*BHS*) is stable also through the 1985/86 period, when again a surge in the M1 aggregate caused instability problems for other previously estimated models. However, the explanation offered by the *BHS* model for the various apparent instability episodes over the whole 1960-1988 period, mainly based on the role of variables capturing long bond yield and interest rate volatility, has been recently challenged under several respects. Boughton (1993) argued that the good performance of the *BHS* model is due more to an extended dynamics than to the introduction of long yield and volatility measures (Hendry and Starr (1993), using encompassing and stability tests, reaffirm the important role played by these variables in determining the fit and stability of the whole model). More importantly, Hess, Jones and Porter (1994) show that the *BHS* model displays serious instability when the sample period is extended to 1993 and attribute this to the excessive weight given to the volatility measure (by construction a backward-looking and slowly reacting variable), leading to an underestimation of the interest rate elasticity of money demand.

be interpreted as reduced forms of a more complex model describing the interrelationships between the monetary base, the money stock, interest rates, income and prices (Gordon (1984)); or they may be affected by measurement errors on the income and interest rate variables (Goodfriend (1985)); or, finally, they may represent reduced forms of an underlying expectations model based on agents' forward-looking behaviour (Cuthbertson and Taylor (1990)). According to all these views (presented in some more detail in the next section), the instability of money demand functions may well be due to shifts in the process generating some of the determinants of money demand, maybe following changes in the policy regime, rather than reflecting variations in the behavioural parameters of the underlying long-term relation.

The instability displayed by empirical dynamic models of money demand has motivated the research aimed at estimating directly the *long-run* relationship between monetary aggregates, income and interest rate variables, using the recently developed cointegration techniques. For the U.S. the results are mixed. On the one hand, B. Friedman (1988a,b) and B. Friedman and Kuttner (1992) view the latest episodes of instability of money demand functions as symptoms of the breakdown of the quantitative relationship between nominal income movements and the growth of monetary aggregates prevailing in earlier decades.[6] Supported by the outcome of various kinds of statistical tests, this view extends also to the broad credit aggregates proposed by B. Friedman himself (1983) as more reliable financial quantities on which targeting policies should be based. On the other hand - following the earlier approach of Meltzer (1963), Chow (1966) and Laidler (1966)- Rasche (1987), Lucas (1988), Hafer and Jansen (1991) and Hoffman and Rasche (1991) found specifications of the long-run function with stable income and interest rate elasticities. In particular, Hoffman and Rasche provide strong evidence of a stable equilibrium money demand function in the post-war period, relating narrowly defined monetary aggregates (the monetary base and M1) to real income, with unitary elasticity, and to nominal interest rates, with an elasticity of -0.5/-0.6 for M1. Moreover, the difference between the actual and the estimated equilibrium level of real balances does not show the high degree of persistence implied by conventional money demand equations based on partial adjustment mechanisms. This finding supports the interpretation of the usually large coefficient on the lagged dependent variable as reflecting the non-stationarity of the series rather than the presence of sizable costs of adjusting money balances. Stock and Watson (1993) investigate the long-run

[6] These results are partly confirmed by Miller (1991), who detects a stable long-run relationship of money balances with their determining variables only for the broader M2 aggregate, but not for M1.

demand for money (M1) using a long sample of annual data (1900-1989), finding a stable function; however, when post-war quarterly data are employed, imprecise estimation of the long-run elasticities occurs. In the U.K. literature, extensive work on the long-run properties of various monetary aggregates has been recently conducted by Hall, Henry and Wilcox (1990). Their main finding is that stable long-run relations between such aggregates and their determinants do exist, provided some variables reflecting financial innovations and broad wealth effects are entered into the estimated (cointegrating) equations. In their comparative study of the stability of long-run money demand in five major countries using postwar quarterly data, Hoffman, Rasche and Tieslau (1995) detect for both the U.S. and the U.K. some episodes of parameter instability over the 1974-1990 period; this instability is eliminated when a (statistically not rejected) unit long-run income elasticity is imposed in estimation.

In general, even when the purpose of the analysis is the direct estimation of the long-run parameters of the money demand function, there may be difficulties in the structural interpretation of estimated coefficients. In fact, when money demand is viewed in the broader context of a system of variables, comprising income, interest rates and inflation, the possibility arises of the existence of multiple long-run relations among the variables. Hypothesis testing on the long-run estimated parameters and examination of the dynamic adjustment of the whole system are necessary for deriving implications for the underlying structural money demand function. A simple example illustrating this point is provided in the next section to motivate the empirical analysis of chapter 3.

3. *Non-structural interpretations of money demand regressions.*

As previously noted, various non-structural interpretations of the estimates from conventionally specified money demand regressions have been put forward in the literature. The present section describes the essence of four such interpretations, starting from the general view of money demand equations as reduced forms of multi-equation systems (section 3.1). The effect of measurement errors in the independent variables on the interpretation of estimated coefficients as structural parameters is then discussed (3.2). Section 3.3 focuses on the comparative stability properties of backward- and forward-looking models when the processes generating the regressors are subject to (maybe policy-induced) shifts. Finally, section 3.4 illustrates the difficulties in making structural inferences on the long-run parameters of the money demand function in the presence of multiple long-run relations linking money balances, income and interest rates.

3.1. Money demand equations as reduced forms.

Conventional partial-adjustment money demand equations have been variously reinterpreted as reduced forms from systems of structural relations. For example, Laidler (1982, 1988) and Gordon (1984) favour the view that conventional money demand equations are really (semi-)reduced forms for the (slowly-adjusting) price level. Accepting this view, the omission of supply-side variables (e.g. oil prices), which can directly affect the price level, may well be one of the causes of instability. Also variations in the degree of sluggishness of the price level may determine changes in the estimated short-run responses of money balances to income and interest rates. However, at least in Laidler's and Gordon's formalizations of this view, the *long-run* money demand elasticities can be estimated correctly. Therefore, even though potentially useful in explaining systematic forecasting errors over periods in which exogenous factors have played a major role in determining the price level (e.g. the oil shocks of the 1970s and early 1980s), such interpretation necessarily ascribes instability of the estimated long-run elasticities to changes in the underlying behavioural parameters.

When other relations linking money, income and interest rates are considered, also the features of the prevailing monetary policy regime are relevant to the interpretation of the estimated coefficients as structural money demand parameters and shifts in the monetary policy rules may induce instability of the estimated money demand equations. The analysis of a stylized equilibrium model of the money market (along the lines set out in Gordon (1984)) may be useful to make this point precise. The model is composed of the following structural (static) money demand and money supply equations (all variables are in logs, except the interest rate):

$$m_t^d = p_t + \alpha_1 y_t - \alpha_2 R_t + \varepsilon_t^d \tag{1}$$

$$m_t^s = \beta_1 B_t + \beta_2 R_t + \varepsilon_t^s \tag{2}$$

where nominal money supply depends on the monetary base B and the interest rate R, with β_1 and β_2 mainly capturing the behaviour of the banking system. The price level (p) and real income (y) are assumed to be exogenously determined, ϵ^d and ϵ^s are demand and supply disturbances respectively, and money demand equals money supply in equilibrium. A unit elasticity of nominal money demand to the price level is assumed. The model is closed by a monetary control rule, assuming different forms according to whether the central bank pursues a monetary base rule or an *interest rate rule*. An example of the latter is given by:

$$R_t = kR_{t-1} + (1-k)\left[R_0 + \delta_1(y_t-\bar{y}) + \delta_2(m_t-m_{t-1})\right] + \varepsilon_t^R \tag{3}$$

Here the current interest rate partially adjusts ($0<k<1$) to a target level set by monetary authorities around the steady-state equilibrium level R_0 (prevailing when $y_t=\bar{y}$ and $m_t=m_{t-1}$), reacting to the rate of monetary growth and to the gap between current and full-employment output $\bar{y}$. Alternatively, an example of a *monetary base rule* is the following: ,

$$B_t = B_{t-1} - \phi_1(y_t-\bar{y}) - \phi_2(p_t-p_{t-1}) \tag{4}$$

where the authorities set the growth rate of B reacting to the inflation rate and to the output gap. Solving the model under the interest rate rule (3) we get the following (reduced form) expression for real money balances:

$$\begin{aligned} m_t-p_t = {} & \frac{\alpha_2\delta_2(1-k)}{1+\alpha_2\delta_2(1-k)}(m_{t-1}-p_{t-1}) + \frac{\alpha_1-\alpha_2\delta_1(1-k)}{1+\alpha_2\delta_2(1-k)}\, y_t - \frac{\alpha_2 k}{1+\alpha_2\delta_2(1-k)}\, R_{t-1} - \\ & - \frac{\alpha_2\delta_2(1-k)}{1+\alpha_2\delta_2(1-k)}\,(p_t-p_{t-1}) - \frac{\alpha_2(1-k)(R_0-\delta_1\bar{y})}{1+\alpha_2\delta_2(1-k)} + \frac{1}{1+\alpha_2\delta_2(1-k)}\,(\varepsilon_t^d-\alpha_2\varepsilon_t^s) \end{aligned} \tag{5}$$

This equation closely resembles conventional empirical "money demand" equations, but now its coefficients clearly depend upon the parameters of the interest rate process, and in particular on the two policy reaction parameters δ_1 and δ_2. Also the long-run "money demand" elasticities derived from (5) differ from their structural values α_1 and α_2. From (5) we have:

$$\mathrm{E}_{(m-p),y} = \alpha_1 - \alpha_2\delta_1(1-k) \quad , \quad \mathrm{E}_{(m-p),R} = -\alpha_2 k \tag{6}$$

and again some parameters of the interest rate process enter the expressions.[7]

On the other hand, if monetary authorities adopt a monetary base rule, the (reduced form) equation for real money balances becomes:

[7] Another notable feature of (5) is the presence of the inflation rate, which does not play any role in the structural money demand equation (1), with a coefficient of opposite sign but of the same magnitude as that on lagged real money balances. This restriction is often tested in the literature (among others by Fair (1987) and McKinnon and Milbourne (1988)) with the aim of distinguishing the hypothesis of partial adjustment of *nominal* money balances (favoured if the restriction is not rejected) from that of partial adjustment of *real* money balances. However, the interpretation of a result in favour of the nominal adjustment alternative may be difficult to sustain since, as shown in the simple example above, such coefficient restriction may quite easily be derived from a model without *any* form of partial adjustment of money holdings in the structural money demand equation.

$$m_t - p_t = \frac{\alpha_2}{\beta_2+\alpha_2}(m_{t-1}-p_{t-1}) + \frac{\alpha_1\beta_2-\alpha_2\beta_1\phi_1}{\beta_2+\alpha_2} y_t - \frac{\alpha_2\beta_2}{\beta_2+\alpha_2} R_{t-1} -$$

$$- \frac{\alpha_2(1+\beta_1\phi_2)}{\beta_2+\alpha_2}(p_t-p_{t-1}) - \frac{\alpha_2\beta_1\phi_1\bar{y}}{\beta_2+\alpha_2} + \frac{1}{\beta_2+\alpha_2}(\beta_2\varepsilon_t^d + \alpha_2\varepsilon_t^s - \alpha_2\varepsilon_{t-1}^s) \tag{7}$$

and the resulting long-run "money demand" elasticities are:

$$\mathrm{E}_{(m-p),y} = \alpha_1 - \frac{\alpha_2\beta_1\phi_1}{\beta_2} \quad , \quad \mathrm{E}_{(m-p),R} = -\alpha_2 \tag{8}$$

Again, the parameters of the monetary rule, as well as the structural parameters of the money supply equation, enter the coefficients of (7) and (one of) the long-run elasticities.

As far as the issue of stability is concerned, the simple example presented, deriving a conventional money demand equation from the minimal departure from a single-equation approach (i.e. a demand/supply model of the money market), can illustrate three main points:

a) the detection of instability may be due to a shift in the conduct of monetary policy from an interest rate to a monetary base rule, determining a shift from (5) to (7) in the estimated equation and a change in the nature and interpretation of long-run solutions;

b) even less dramatic monetary policy changes, as the modification of the policy parameters capturing the degree of reaction of the target variable to the state of the economy (summarized by the inflation rate, the output gap or the rate of money growth), may generate instability of the estimated equations;

c) finally, if a monetary base rule is followed, also changes in the behaviour of the banking system or other supply-side factors may determine structural instability.

All these factors may be responsible for structural breaks -detected, for example, by recursive stability tests- and variability of the long-run solutions, with no change in the underlying structural parameters of the money demand function, leading to incorrect inferences about modifications of agents' behavioural characteristics.

3.2. Measurement errors in the independent variables.

Further reasons of caution in interpreting the detected parameter instability as structural are provided by an alternative view of conventional partial adjustment money demand equations originally proposed by Goodfriend (1985). The central point is that if the determinants of money demand (namely income and interest rates) are only measured with a stochastic error, then a money demand equation in a partial adjustment form can be derived, but its income and interest rate coefficients will also depend on the parameters of the generating process of the appropriate measures of the variables and on the magnitude of the measurement errors.

In order to highlight the potential role of policy regime shifts also under this alternative view of money demand regressions, let us consider the effect of stochastic errors only in the measure of the interest rate in a single-equation context. The structural money demand function is then reformulated as follows:

$$m_t = p_t + y_t - \alpha_2 R_t^* + \varepsilon_t^m \tag{9}$$

Here unit elasticities to income and the price level have been assumed for simplicity, and R^* is the measure of the opportunity cost of holding money which is relevant in determining agents' behaviour. R^* evolves through time according to a simple first-order autoregressive process:

$$R_t^* = \theta_0 + \theta_1 R_{t-1}^* + \varepsilon_t^R \qquad\qquad 0 < \theta_1 < 1 \tag{10}$$

where ϵ^R is a white noise stochastic element with variance σ^2_R. The relationship between R^* and the interest rate variable actually included in the estimated regression is:

$$R_t = R_t^* + e_t \tag{11}$$

with e denoting the white noise measurement error, independent of ϵ^m and ϵ^R, and with variance σ^2_e. The conventional (partial adjustment) money demand regression, including R and with correctly imposed unit elasticities to y and p, is of the following form:

$$m_t - p_t - y_t = b_0 + b_2 R_t + b_3(m_{t-1} - p_{t-1} - y_{t-1}) + u_t \tag{12}$$

Given the underlying model described by (9), (10), and (11), the estimates of b_2 and b_3 will be:

$$b_2 = -\frac{\sigma_R^2}{\sigma_R^2 + \sigma_e^2}\alpha_2 \quad (-\alpha_2 < b_2 < 0) \quad \textit{and} \quad b_3 = \frac{\theta_1 \sigma_e^2}{\sigma_R^2 + \sigma_e^2} \quad (0 < b_3 < 1) \tag{13}$$

The presence of a measurement error in the interest rate ($\sigma^2_e > 0$) determines a downward bias in the estimate of α_2 and attributes significance to the lagged (inverse) velocity term, which is now able to predict $(m\text{-}p\text{-}y)_t$ in the presence of autocorrelation in the process generating R^*. Also the parameters in (10) affect the estimated coefficients.[8] The

[8] In the case considered above, only b_3 depends on θ_1, but the inclusion of an additional measurement error in y would make also b_2 dependent on the parameters of the processes generating the relevant variables R^* and y^*.

estimated long-run elasticity of real money balances to the measured interest rate is:

$$E_{(m-p-y),R} = -\alpha_2 \frac{\sigma_R^2}{\sigma_R^2 + (1-\theta_1)\sigma_\epsilon^2} \tag{14}$$

being a function of σ^2_R, σ^2_ϵ and θ_1, in addition to the structural parameter α_2. Again, detected changes of the estimate of such elasticity or more general instability problems can occur in the face of shifts in the process generating R^*, with no change in structural parameters. If monetary policy is able to affect the time series properties of the relevant measure of the opportunity cost of money holdings, then monetary policy regime shifts may again be responsible for the detected instability even in this more limited single-equation context.

The measurement error view of conventional money demand regressions has interesting implications for more general theoretical issues, as pointed out by Laidler (1985),[9] but its empirical relevance can only be assessed if precise hypotheses about the relationships between the true determinants of money demand and the proxies commonly used in empirical work are formulated. Under this respect, the deficiencies of measured GNP as the appropriate transactions variable (due, for instance, to the existence of the underground economy or to the neglect of financial transactions) and of average or end-of-period interest rates as measures of the theoretically appropriate opportunity cost of holding money, are widely cited. Furthermore, the fact that actual values of income and interest rates measures are used in the estimates when agents base their decisions on expectations of these variables may be an additional cause of measurement error, leading to potential instability of the equation.[10]

3.3. Money demand instability and the role of expectations.

The potential role of expectations in improving the stability of money demand functions has been the focus of some recent research aimed at constructing models explicitly

[9] In particular, this view explains the significance of lagged money balances in the estimated equations not as reflecting slow money holdings adjustment or (as in Laidler's (1982) interpretation) price level stickiness, but as due to the fact that lagged money helps predict current money balances in the presence of autocorrelation in the process generating the true income and interest rate variables. This implies that the economy always operates on its long-run aggregate demand for money function, potentially reconciling the equilibrium (neo-classical) view of the market mechanism with the available empirical evidence.

[10] Assuming measurement error only in the income variable, Taylor (1994) derives testable implications of the Goodfriend hypothesis for the dynamic specification of money demand regressions. When applied to U.S. data for a period in which the stability of conventional equations is uncontroversial (1952-1972), the test yields a strong rejection of the measurement error hypothesis.

derived from agents' forward-looking behaviour, as the multi-period costs of adjustment models of Cuthbertson and Taylor (1987, 1990) and Cuthbertson (1988), and the dynamic rational expectations framework without adjustment costs proposed by Dutkowsky and Foote (1988, 1992). The essential point of this strand of literature is that if agents adopt forward-looking behaviour in their decisions on money holdings, then the estimated parameters of traditional money demand functions, neglecting the role of expectations, represent convolutions of *deep* structural parameters and parameters describing the process generating expectations. Consequently, the empirically detected money demand instability could be wrongly attributed to shifts in the structural coefficients of the (long-run) function, the actual cause being the instability of the expectations generating process. In the face of the success of a class of feedback models, mainly based upon the error-correction mechanism, in explaining the demand for money even in times of instability (e.g. Hendry (1985) and Hendry and Ericsson (1991) for the U.K., Rose (1985) and Baba, Hendry and Starr (1992) for the U.S.), the consequences for such specifications of neglecting expectations formation have been analyzed by Kelly (1985), Cuthbertson (1986a, 1991), Taylor (1987), Hendry and Neale (1988), Hendry (1988) and Favero and Hendry (1992). Two main conclusions, relevant to our discussion, have been reached.

The first -already referred to above- concerns the nature of the short- and long-run elasticities estimated from conventional distributed lag functions and may be presented with the aid of the following simple expectations model describing the behaviour of real money balances:

$$(m-p)_t = \alpha_0 + \alpha_1(L)y_t + \alpha_2 \mathrm{E}(R_t \mid \Omega_{t-1}) + \alpha_3(L)R_{t-1} + \alpha_4(L)(m-p)_{t-1} + \varepsilon_t$$
$$\alpha_1(L) = \sum_{h=1}^{H} \alpha_{1h}L^{h-1} \quad , \quad \alpha_3(L) = \sum_{j=1}^{J} \alpha_{3j}L^{j-1} \quad , \quad \alpha_4(L) = \sum_{i=1}^{I} \alpha_{4i}L^{i-1} \tag{15}$$

with the information set available to agents defined as $\Omega_{t-1} = \{(m-p)_{t-1}, (m-p)_{t-2}, \ldots, y_{t-1}, y_{t-2}, \ldots, R_{t-1}, R_{t-2}, \ldots\}$. The covariance stationary process generating R_t is assumed to be:

$$R_t = \theta_0 + \theta_1(L)R_{t-1} + v_t$$
$$\theta_1(L) = \sum_{k=1}^{K} \theta_{1k}L^{k-1} \quad , \quad \theta_1(1) < 1 \tag{16}$$

with v_t and ϵ_t being uncorrelated white noise disturbances. The condition $K > J$ is imposed in order to allow identification of the parameters α_{3j}. Using (16) to substitute for expectations in (15) we obtain:

$$(m-p)_t = (\alpha_0+\alpha_2\theta_0) + [\alpha_2\theta_1(L)+\alpha_3(L)]R_{t-1} + \alpha_1(L)y_t + \alpha_4(L)(m-p)_{t-1} + \varepsilon_t$$
$$= \quad \mu_0 \quad + \quad \mu(L)R_{t-1} \quad + \alpha_1(L)y_t + \alpha_4(L)(m-p)_{t-1} + \varepsilon_t \qquad (17)$$

where

$$\mu_i = \begin{cases} \alpha_2\theta_{1i} + \alpha_{3i} & for \quad i = 1,...J \\ \alpha_2\theta_{1i} & for \quad i = J+1,...K \end{cases} \qquad (18)$$

Kelly (1985) argues that estimation of (17) under the assumption of "exogeneity" of R leads to an estimate of the long-run solution which cannot be interpreted as the true behavioural elasticity, but depends on the parameters of the data generating process of the (rationally) expected variable. In fact, the long-run elasticities of real money balances with respect to y and R are:

$$E_{(m-p),y} = \frac{\alpha_1(1)}{1-\alpha_4(1)} \quad , \quad E_{(m-p),R} = \frac{\alpha_2\theta_1(1)+\alpha_3(1)}{1-\alpha_4(1)} \qquad (19)$$

yielding a long-run elasticity of *(m-p)* with respect to R which is a function of the parameters θ's (since $\theta(1)<1$). The implication is that if expectations play an important role in behavioural relationships, then inferences about the true steady-state behavioural parameters based only on the long-run equilibrium solutions of conventional distributed lag equations such as (17) can be highly misleading.

Reinterpreting Kelly's analysis, Hendry and Neale (1988) argue that the above results derive from an incorrect exogeneity assumption about R. Specifically, since from (17) α_2 and α_{3j} cannot be estimated without knowledge of the θ's in the interest rate process, then R_t is not weakly exogenous for such parameters α's.[11] This implies that not only the inferences about the long-run solutions but also those about short-run responses of real money balances to the interest rate are incorrect, since the coefficients in $\mu(L)$ are different from those in $\alpha_3(L)$. However, if *(m-p)* and R are integrated of order one (*I(1)*) and cointegrated, 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: under rational expectations, actual and expected values differ only by a stationary (*I(0)*) expectational error which by its nature does not affect the estimated long-run (cointegrating) relation.

[11] Various concepts of exogeneity (*weak, strong* and *super*exogeneity) are discussed by Engle, Hendry and Richard (1983) and associated, as necessary requirements, to the different utilizations of a model (respectively, hypothesis testing, forecasting and policy analysis).

The second main point of interest concerns the possibility of discriminating empirically between structural feedback models and feedback models which are reduced forms of forward-looking structural models by means of structural stability and (variance) encompassing tests (Hendry (1988, 1994)). To illustrate this possibility consider the following two alternative behavioural hypotheses:

$$\textit{feedback} \quad H_c: \quad (m-p)_t = \alpha' x_t + \varepsilon_t \qquad \mathrm{E}(x_t \varepsilon_t) = 0 \tag{20}$$

$$\textit{forward-looking} \quad H_e: \quad (m-p)_t = \delta' \mathrm{E}(x_t \mid z_{t-1}) + \phi_t \tag{21}$$

$$\textit{and} \quad x_t = \pi_t z_{t-1} + w_t \qquad \mathrm{E}\left(z_{t-1} w_t'\right) = 0 \ , \ \mathrm{E}\left(w_t w_t'\right) = \Sigma_t \tag{22}$$

where $x_t = (y_t, R_t)$. The feedback hypothesis in (20) implies that agents act on the basis of the *observed* current values of income and the interest rate. Therefore, given x_t, the variables in z_{t-1} are irrelevant for the determination of $(m-p)_t$. On the other hand, the forward-looking hypothesis, consisting of the structural model (21) and the marginal model for x_t specified in (22) with time-varying parameters, implies the invalidity of conditioning on x_t.

Each hypothesis in turn is now assumed to be the correct characterization of the data generating process (*DGP*) of $(m-p)_t$ and x_t and the corresponding implications in terms of parameter stability and error variance are derived:

(*i*) when the feedback model (20) is a correct representation of the *DGP*, the estimated form generated by the forward-looking alternative in (21), using (22), is:

$$\begin{aligned} (m-p)_t &= \alpha' \pi_t z_{t-1} + (\varepsilon_t + \alpha' w_t) \\ &= \Theta_t z_{t-1} + \quad v_t \\ \textit{with} \quad \sigma^2_{vt} &= \sigma^2_\varepsilon + \alpha' \Sigma_t \alpha \ > \ \sigma^2_\varepsilon \end{aligned} \tag{23}$$

The parameters in (23) vary, following changes in the process generating x_t, whereas those of the feedback model (20) are (by assumption) constant and the standard error of the regression σ_v is larger than that of the feedback model, σ_ε, and may vary over time;

(*ii*) when the forward-looking model (21) is a correct representation of the *DGP*, the estimated form generated by the feedback alternative (20) is:

$$\begin{aligned} (m-p)_t &= \delta' x_t + (\phi_t - \delta' w_t) \\ &= \delta' x_t + \quad u_t \\ \textit{with} \quad \sigma^2_{ut} &= \sigma^2_\phi \ \delta' \Sigma_t \delta \ - \ 2\delta' \eta_t \end{aligned} \tag{24}$$

where $\eta_t = E(w_t\phi_t)$. In this case the conditional model (24) cannot be constant if π_t is sufficiently variable, since

$$\mathrm{E}\big((m-p)_t \mid x_t\big) = \delta_t^{*\prime} x_t \tag{25}$$

$$\begin{aligned} \textit{with} \quad \delta_t^* &= \left[\mathrm{E}(x_t x_t')\right]^{-1} \mathrm{E}\big(x_t (m-p)_t\big) \\ &= \left[\pi_t M_{zz} \pi_t' + \Sigma_t\right]^{-1} \left[\pi_t M_{zz} \pi_t' \delta + \eta_t\right] \\ &= \delta + \left[\pi_t M_{zz} \pi_t' + \Sigma_t\right]^{-1} \left[\eta_t - \Sigma_t \delta\right] \qquad M_{zz} = \mathrm{E}(z_{t-1} z_{t-1}') \end{aligned} \tag{26}$$

which displays non-constancy due to the time-varying nature of the process generating x_t. The outcome of the comparison of the standard error of the regression σ_u with that of the forward-looking model σ_ϕ depends on the relative magnitudes of the two terms involving Σ_t and η_t. Obviously, if $E(w_t\phi_t) = 0$, we get $\sigma^2_u > \sigma^2_\phi$, reversing the error variance ranking found under (*i*).

The foregoing analysis gives rise to two main conclusions which are relevant to the issue of distinguishing empirically between the two rival hypotheses:

a) given the implications of the two alternative models for the standard error of the regression, encompassing tests comparing the error variances of the two estimated models *may* be able to discriminate between them (this is the case, for example, if $E(w_t\phi_t) = 0$);

b) more importantly, given the above implications in terms of parameter stability, if the conditional model $(m\text{-}p)_t = \alpha' x_t + \epsilon_t$ has α constant but the marginal model for x_t, $x_t = \pi_t z_{t-1} + w_t$, has π_t non-constant, then the interpretation of the feedback model as derived from an underlying forward-looking structure is *not* sustainable. In other words, any non-constancy in the process generating x_t and used by agents in forming expectations must be reflected in the non-constancy of the conditional model if the forward-looking hypothesis represents the correct characterization of the DGP. The above statement on stability can be shown to apply also when z_{t-1} represents only part of the agents' information set. In this practically relevant case the actual data generating process for x_t is:

$$x_t = \pi_1 z_{t-1} + \pi_2 z_{t-1}^* + a_t \qquad \mathrm{E}(a_t a_t') = \Lambda \tag{27}$$

and agents form their expectations accordingly on the basis of the complete information set $I_{t-1} = \{z_{t-1},\ z^*_{t-1}\}$, whereas only the variables in z_{t-1} are included in the estimated marginal model for x_t. It could then appear that the non-constancy of the estimated marginal model, which, together with the constancy of the conditional model, leads to the rejection of the forward-looking alternative, is only due to an incorrect formulation of the process generating x_t, the true process in (27) having constant parameters π_1 and π_2.

However, the detected non-constancy of π_t must be due to a non-constant relationship between the included (z_{t-1}) and the omitted (z^*_{t-1}) variables in the model for x_t. If such a relationship can be written as:

$$z^*_{t-1} = A_t z_{t-1} + b_t \qquad\qquad E(b_t b_t') = B_t \tag{28}$$

then the estimated parameters of the now incomplete marginal model (22) are given by:

$$\pi_t = \pi_1 + \pi_2 A_t \quad , \quad \Sigma_t = \Lambda + \pi_2' B_t \pi_2 \tag{29}$$

This non-constant relationship between z and z^* implies that even if the correct model (27) is used, the regression of $(m\text{-}p)_t$ on x_t cannot yield constant parameters, since in this case the π matrix in (26) is constant but the analogue of the moment matrix M_{zz} cannot be. Therefore, if only a subset of the complete information set available to agents is used in estimation, the proposition stated above is still valid: the joint constancy of the conditional model and the non-constancy of the marginal model imply that the interpretation of the feedback specification as the reduced form of an expectational structure is *not* acceptable.

Overall, the detected non-constancy of the parameters of the estimated model for x_t may be due either to the non-constancy of the data generating process of x_t or to a non-constant relationship between the included and the omitted variables when only a subset of the information available to agents is used in the analysis. In both cases, however, if the forward-looking alternative is correct, the estimated parameters of the conditional model must display non-constancy, allowing one to discriminate empirically between the two structures. To this aim, the analysis of the stability performance of empirical specifications of the money demand function by means of recursive stability tests seems extremely useful. However, various considerations suggest caution in interpreting the results of stability tests on feedback and marginal models.

First, as shown by Favero and Hendry (1992) using Monte Carlo simulations, constancy (Chow) tests performed on an invalid conditional model have low power in detecting instability in the face of shifts in the parameters (mean and variance) of the marginal models. On the contrary, conventional model mis-specification (i.e. the omission of relevant variables) may be responsible for structural instability, readily detected by constancy tests, when the processes generating the omitted variables are subjected to shifts. These results imply that the Lucas (1976) critique of conventional feedback models, though theoretically valid, may be of limited *empirical* relevance. Moreover, they call for caution in interpreting the instability of conventional feedback models (even when the marginal models for the regressors display instability) as immediate evidence in favour of a forward-looking alternative.

Second, the stability implications of feedback and forward-looking models illustrated above show that if a backward-looking specification of the demand for money function is found which is sufficiently stable, whereas the processes generating the regressors are highly unstable, a structural expectations model can be rejected, even without having to specify and estimate the forward-looking alternative. However, as noted by Cuthbertson (1991), *in finite samples* it may be relatively easy to find instability in the marginal models for the regressors, searching over various arbitrary information sets, even if the true process generating x in equation (22) above has time-invariant parameters, and the *population* moment matrix of the z variables (the generalization of M_{zz} in (26), including z and z^*) is time-invariant as well. Hence, a result of instability of the estimated marginal model, together with parameter constancy of the feedback specification, might simply be due to the fact that the econometrician is using an incorrect marginal model in a finite sample, instead of showing the inadequacy of a forward-looking alternative.[12] On this basis, Cuthbertson favours the implementation of stability tests directly on the estimated structural forward-looking model, to be compared with the stability performance of the feedback alternative.

The above considerations point towards the analysis of specific historical episodes, where clear changes in the time-series processes of the determinants of money demand can be identified and related to specific causes, as a more powerful way of assessing the role of expectations in determining the demand for money balances. In particular, monetary policy regime shifts are potential candidates, since they are often readily reflected in the characteristics of interest rate behaviour. Indeed, if the marginal model for interest rates shows a structural break at the relevant (policy regime shift) dates, such result may not so easily be due only to finite sample variability, but may well indicate a policy-induced radical change in the time-series properties of interest rates. Then, the fact that a feedback model for money demand shows similar breaks at the same dates may strongly suggest that conditioning on the interest rates variables is invalid and that a forward-looking alternative may be more appropriate.

The empirical investigation of the next chapter follows these lines, analyzing the stability of a feedback model of the Italian money demand in the face of a clear change in the prevailing monetary policy regime and comparing the results with the stability performance of a forward-looking alternative specification.

[12] It should be noted that Cuthbertson's point is not about the theoretical validity of the conclusions stated in the preceding section, but it concerns their *practical* relevance, since the analyst is always likely to formulate an incorrect marginal model *and* is forced to work with finite samples.

3.4. Multiple long-run relations and single-equation models.

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. In fact, the likely existence of simultaneity between money holdings and their determinants may require an estimation method based on instrumental variables. 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. Cointegration analysis[13] provides formal procedures to detect the existence of multiple long-run relations among the variables. Even when simultaneity is not a relevant problem, overlooking the presence of more than one cointegrating relation may lead to serious misinterpretations of the long-run properties of agents' behaviour and also to mis-specifications of the short-run dynamic adjustment towards equilibrium.

To illustrate this point, consider four variables (lowercase 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:

$$m - p = \alpha y \tag{30}$$

$$R^m = \gamma R^b \tag{31}$$

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:

$$\Delta(m-p)_t = a_1 \Delta(m-p)_{t-1} - a_2[(m-p)-\alpha y]_{t-1} + a_3(R^m - \gamma R^b)_{t-1} + u_{1t} \tag{32}$$

$$\Delta y_t = b_1 \Delta y_{t-1} + b_2[(m-p) - \alpha y]_{t-1} + u_{2t} \tag{33}$$

$$\Delta R_t^m = c_1 \Delta R_{t-1}^m - c_2(R^m - \gamma R^b)_{t-1} + u_{3t} \tag{34}$$

$$\Delta R_t^b = d_1 \Delta R_{t-1} + u_{4t} \tag{35}$$

Both money balances and expenditure react to past deviations of money demand from the

[13] Among the most relevant contributions to this literature are Engle and Granger (1987), Johansen (1988, 1991), Johansen and Juselius (1990, 1992, 1994), Banerjee, Dolado, Galbraith and Hendry (1993). Campbell and Perron (1991), Muscatelli and Hurn (1992) and Ericsson (1992) survey the field and provide extensive bibliographies.

equilibrium (long-run) relation (30). Also the interest rate on money displays error-correcting behaviour, since the relevant disequilibrium term enters equation (34). 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 (32)-(35) 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 (32), is the following:

$$\Delta(m-p)_t = \delta_1\Delta(m-p)_{t-1} - \delta_2(m-p)_{t-1} + \delta_3 y_{t-1} + \delta_4 R^m_{t-1} - \delta_5 R^b_{t-1} + \epsilon_t \tag{36}$$

The estimated long-run solution, obtained from the terms in levels in (36), 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 $(m\text{-}p)\text{-}\alpha y$ and $R^m\text{-}\gamma R^b$ in (33) and (34) imposes (testable) cross-equation restrictions on the system parameters. These restrictions, either implied by some economic theory or suggested by unrestricted estimation of (32)-(35), 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.

In chapter 3 a multivariate approach to the specification of money demand is adopted and applied to Italian data for the 1980s and early 1990s. Multiple long-run relations among the variables analyzed are found and structural hypotheses on these are formally tested. A complete simultaneous system is then estimated with a dynamics consistent with the proposed economic interpretation of the long-run equilibrium relationships. Finally, the results obtained from the system estimation are compared with the long-run money demand equation derived from a single-equation approach, illustrating the difficulties in the economic interpretation of the estimated coefficients.

Chapter 2

Money demand instability, expectations and policy regimes: an application to Italy (1964-1986).

1. *Introduction*

The empirical analysis of the present chapter compares feedback and forward-looking models of the demand for money using Italian data for the period 1964-1986.

In recent years, when applied to U.S. or U.K. data, both classes of models seemed capable of yielding satisfactory characterizations of money demand behaviour. On the one hand, the class of feedback models based upon the error-correction mechanism has proved successful in explaining the demand for money even in times of high instability (e.g. Hendry (1988) and Baba, Hendry and Starr (1992)). Recent developments in the theory of cointegration have provided a more rigorous statistical background to the error-correction approach and offered a relatively simple empirical specification strategy to model the long-run equilibrium relation and the short-run dynamics between economic variables. On the other hand, the role of expectations on the future evolution of the determinants of money holdings is the main feature of the multi-period cost-of-adjustment models, successfully applied to both the U.S. and the U.K. by Cuthbertson and Taylor (1987, 1990).

Comparative assessments of the performance of these two classes of models for the U.K. have been recently provided by Hendry (1988), Muscatelli (1989) and Cuthbertson and Taylor (1992), using different evaluation methods. Muscatelli compares the two models on the basis of several model selection criteria and of the results of variance encompassing tests. His conclusions favour the feedback model, specified by means of a general to specific strategy. Hendry contrasts his (1985) feedback equation with the forward-looking model of Cuthbertson (1988), providing one application of the stability analysis theoretically illustrated in the preceding chapter. The non-constancy of the autoregressive processes used by Cuthbertson to generate expectations, together with the remarkable stability of the feedback specification, is viewed as strong evidence against Cuthbertson's interpretation of the feedback model as an approximation to an underlying expectational structure.

In the present chapter we apply this kind of analysis to the behaviour of Italian money demand. The case of Italy seems interesting mainly because the analysis of the time-series behaviour of the interest rates entering conventional money demand equations as conditioning variables shows a clear structural break corresponding to a change in monetary policy procedures (namely the abandonment of the stabilization of bond prices) in 1970. Moreover, in the mid-'70s, the processes generating inflation and, again, interest rates, display marked structural breaks.

Attention is therefore focused on such episodes, which may provide a meaningful test of the two alternative models. The comparison is conducted in terms of the relative stability performance for two main reasons. First, the criteria used in Muscatelli (1989) are those with respect to which the feedback specification is selected and therefore seem to unduly favour one of the competing models. Second, the estimation method we chose for the forward-looking model determines by design an increase in the standard error of the estimated equation which makes it inappropriate a comparison on the basis of encompassing tests. In our case, we think that a structural stability analysis is a much more compelling test of the two models.

The chapter is organized as follows. In section 2, after a brief description of the econometric methodology, a feedback model for Italian money demand is specified and a stability analysis is performed. Given the detection of several structural breaks in the final feedback specification, equations for the processes generating the regressors are estimated in section 3. Their stability properties are also assessed and formally related to the detected pattern of instability. Section 4 is devoted to a discussion of the main estimation methods available for forward-looking models and to the specification of an alternative money demand model; again the stability of the model is assessed. Section 5 briefly concludes.

2. *A feedback model for the Italian money demand.*

2.1. The econometric specification of a feedback model.

Formally, a *feedback model* can be defined as a simplified representation of the joint probability of all sample data on both endogenous and exogenous variables (the *Data Generating Process*, *DGP*, of the variables), after *marginalization* with respect to those variables that are unimportant to the determination of the series of interest, and *conditioning* of the endogenous variables on the set of weekly exogenous regressors. It includes only observed variables and is not based on the explicit modelling of expectations. In the terminology introduced by Hendry and Richard (1983) and Gilbert (1986), a satisfactory feedback model should be a *congruent* representation of the data, displaying several desirable properties: data admissibility, consistency with theory, weak exogeneity of the set of regressors, parameter constancy, data coherency (i.e. the requirement that the residuals generated by the model are true innovations with respect to the available information) and encompassing of a wide range of rival models.

On practical grounds, the various strategies that have been formulated in order to obtain congruent empirical feedback models share several common features. Firstly, 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, giving rise to a maybe complex short-run dynamics around the long-run equilibrium. Therefore, to be a congruent representation of the data, a feedback model must capture both the equilibrium relation and the shape of the short-run dynamics of the variables under study. Moreover, all specification strategies require the final model to be a *balanced* representation of the data, in the sense that the statistical properties of the dependent and explanatory variables must be consistent. In particular, in order to apply classical asymptotic results, stationary variables are needed. Since most economic series are non-stationary, balanced relations between stationary variables can be achieved by appropriate differentiation or by considering cointegrating vectors, i.e. stationary linear combinations of non-stationary variables.

The econometric method applied in the empirical analysis of this section aims at a simultaneous specification of the long-run equilibrium relation and the short-run dynamics, by means of the *general to specific* modelling strategy developed and implemented by D. Hendry in a series of papers (for example, Hendry (1985, 1987), and Baba, Hendry and Starr (1992)). The basic idea underlying this methodology is to derive the final specification from a general *baseline* unrestricted dynamic model through several steps of reduction and reparameterization of the included variables. This process involves a loss of information, whose relevance must be assessed by testing procedures designed to check whether the model is a congruent representation of the *DGP*. In particular, the error term must be a true innovation, being unpredictable on the basis of the available information, and the regressors must satisfy the exogeneity requirement for the relevant parameters which is appropriate for the proposed use of the model (hypothesis testing, forecasting or policy analysis).

The baseline model includes the variables that economic theory considers relevant to the problem at hand and contains unrestricted dynamics, in the form of long lags of both the dependent and the independent variables. The generality of the model reflects the belief that theory can only suggest which variables are likely to enter the long-run equilibrium relation, but only the data can determine the shape of the short-run dynamics. Notwithstanding its generality, the baseline model is itself the outcome of some reduction, since there may exist variables included in the DGP but omitted from the chosen general model. Therefore, diagnostic checking procedures designed to test the relevance of the lost information have to be conducted also on the baseline model. Further reductions and reparameterizations can then be implemented by imposing all the restrictions suggested by the data in the form both of exclusion restrictions and of transformations on the levels of the variables, in order to obtain near orthogonal regressors with meaningful economic interpretations (e.g. differences or error-correction terms).

In the context of money demand modelling, a plausible baseline model could take the following form:

$$m_t = c + \sum_{i=1}^{M} \alpha_i \, m_{t-i} + \sum_{i=0}^{M} \beta_{1i} \, p_{t-i} + \sum_{i=0}^{M} \beta_{2i} \, y_{t-i} + \sum_{i=0}^{M} \beta_{3i} \, R_{t-i} + u_t \tag{1}$$

where c is a constant, m, p, and y are the logarithms of the chosen monetary aggregate, the price level and real income, R is a measure of the opportunity cost of holding money balances, and M is the maximum lag (e.g. set at 5 if quarterly data are used). Diagnostic checking on (1) will ensure that u_t is a true innovation, i.e. $E(u_t | I_t)=0$, with $I_t=\{m_{t-1},\ldots,p_t, p_{t-1},\ldots,y_t,y_{t-1},\ldots,R_t,R_{t-1},\ldots\}$. The process of reduction, reparameterization and testing on (1) may lead to the following typical final specification (similar equations can be found in Hendry and Mizon (1978), Hendry (1985) and Rose (1985)):

$$\Delta(m-p)_t = c + \delta_0 \Delta(m-p)_{t-1} + \delta_1 \Delta y_t + \delta_2 (m_{t-1} - p_{t-1} - y_{t-1} + kR_{t-1}) + \delta_3 \Delta p_t + \delta_4 \Delta R_t + \varepsilon_t \tag{2}$$

where, again, diagnostic tests ensure that $E(\epsilon_t \mid I_t) = 0$ and that the model satisfies the other main requirements for congruency with the data. In (2), several restrictions on the baseline model are imposed and tested. Specifically, the following set of exclusion restrictions:

$$\begin{array}{llll} \alpha_i = 0 & \beta_{1i} = 0 & \quad \textit{for} \ \ i>2 \\ \beta_{2i} = 0 & \beta_{3i} = 0 & \quad \textit{for} \ \ i>1 \end{array} \tag{3}$$

and the additional three linear restrictions:

$$\begin{array}{l} \alpha_2 = -\beta_{12} \\ \alpha_1 + \beta_{10} + \beta_{11} = 1 \\ \beta_{10} + \beta_{11} + \beta_{12} = \beta_{20} + \beta_{21} \end{array} \tag{4}$$

must not be rejected. In order for model (2) to represent a balanced money demand equation, the statistical properties of the included variables must be consistent. If, for example, real money balances *(m-p)*, real income, the price level and the interest rate are all integrated of order one (*I(1)*) and cointegrated, with cointegrating vector *(1,-1,-1,k)*, model (2) is a balanced equation, involving only stationary (*I(0)*) variables, being either first differences of *I(1)* series or stationary cointegrating relations $(m_{t-1}-p_{t-1}-y_{t-1}+kR_{t-1})$.

The economic interpretation of empirical models of this class is often based on agents following purely feedback rules of behaviour, reacting to observed (current and lagged) variables when deciding current values of their choice variable. Models of this kind may then represent simplified rules-of-thumb which agents may follow in complex environments (Hendry (1988)). Under this interpretation, if the coefficient on the term in levels (δ_2 in (2)) is negative, it can be argued that this term gives the long-run relation between the set of variables and agents, when deciding the current value of money balances, react to past deviations from equilibrium in such a way that the change in money holdings 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*) for this class of models.[1]

The absence of any role for expectations in the above interpretation of error-correction models has often be regarded as a major drawback, implying sub-optimality of agents' behaviour. However, an error-correction formulation, though typical of a feedback model, may nevertheless reflect, under certain conditions, optimal responses of forward-

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[1] The first empirical model based on the error-correction mechanism is Sargan's (1964) model of UK wage determination. Davidson, Hendry, Srba and Yeo (1978) successfully applied the *ECM* specification to the UK consumption function. The historical evolution of the *ECM* concept and its applications are surveyed by Alogoskoufis and Smith (1991); Hendry, Muellbauer and Murphy (1990) provide a recent reassessment of the whole econometric methodology underlying the *ECM* approach in the context of the UK consumption function.

looking agents in a dynamic environment (Nickell (1985), Dolado (1987)). This point is usually made in the context of the standard intertemporal quadratic adjustment cost model (as developed, for example, by Sargent (1978, 1989)), in which agents are supposed to make a sequence of decisions $\{m_{t+i}\}$ in order to reach the target $\{m_{t+i}^*\}$, with the objective of minimizing the expected present value of a quadratic loss function, incorporating costs of adjusting m to m^*. The representative agent's problem is usually expressed in the following form:

$$\min_{m_{t+i}} \ \mathrm{E}_t \sum_{i=0}^{\infty} \phi^i \left[\, c\,(m_{t+i} - m_{t+i}^*)^2 + (m_{t+i} - m_{t+i-1})^2 \, \right] \quad , \qquad (5)$$

where the loss function incorporates both the cost of being out of equilibrium and the cost of adjusting the control variable, with the positive constant c measuring their relative importance, and ϕ denoting the constant discount factor $(0<\phi<1)$. The first order condition (Euler equation) for this problem is, at time t:

$$\phi E_t m_{t+1} - (1 + \phi + c) m_t + m_{t-1} = - \, c m_t^* \qquad (6)$$

The difference equation (6) may be solved forward by standard factorization methods, since the roots of the characteristic polynomial associated with (6) are both positive and lie on either side of unity. The solution takes then the following form, with μ denoting the stable root:

$$m_t = \mu m_{t-1} + (1-\mu)(1-\mu\phi)\sum_{i=0}^{\infty} (\mu\phi)^i \mathrm{E}_t m_{t+i}^* \qquad (7)$$

Now, a specification of the process generating m^* is needed in order to rewrite (7) in terms of observable variables. If m^* is described by an *ARI(1,1)* stochastic process, being:

$$\Delta m_t^* = \beta \Delta m_{t-1}^* + v_t \qquad (\,|\beta|\,<1) \qquad (8)$$

the decision rule followed by agents takes the following form:

$$\Delta m_t = \frac{1-\mu}{1-\mu\phi\beta} \Delta m_t^* - (1-\mu)(m_{t-1} - m_{t-1}^*) \qquad (9)$$

Equation (9) displays the error-correction mechanism, since the term $(m_{t-1}-m^*_{t-1})$ measures past deviations from target and enters the equation with a negative coefficient.[2] Of course, this result depends crucially on the specific stochastic process for the target variable (or its determinants), but it is of particular interest since the process in (8) seems to describe

[2] Richer dynamics in the form of lags of Δm^*, making (9) more similar to commonly estimated equations, are obtained if m^* follows an *ARI(n,1)* process, with $n>1$.

satisfactorily the behaviour of many economic variables. Further insight on the interpretation of error-correction terms in estimated equations is provided by Dolado (1987), who rewrites the Euler equation (6) as:

$$\Delta m_{t+1} = \frac{1}{\phi}\Delta m_t + \frac{c}{\phi}(m_t - m_t^*) + (\Delta m_{t+1} - \mathrm{E}_t \Delta m_{t+1}) \tag{10}$$

Also (10), where the last term is a true innovation, displays a term representing past deviations from target. However, an error-correction behavioural interpretation is not allowed since the coefficient c/ϕ is positive, implying that a level of m_t greater than m_t^* has a positive impact on the rate of growth of m, therefore amplifying the deviation from target.

Taken together, the above considerations suggest that the presence of terms in lagged levels of the variables in apparently feedback equations may not be interpretable as evidence for rule-of-thumb behaviour, with agents reacting to past deviations from equilibrium, but may well be derived as part of the decision rule followed by fully optimizing, forward-looking individuals. Discrimination between the two alternatives cannot be simply based on the sign of the coefficient of the lagged level term since, under reasonable assumptions, a negative coefficient may be generated by forward-looking behaviour.[3] As outlined in the previous chapter, a comparative stability analysis involving marginal models for the regressors in the feedback specification may be more informative on which alternative is a more adequate description of the data.

The recent literature on cointegration has provided formal statistical foundations for error-correction modelling, showing that if some *I(1)* series are cointegrated, so that a linear combination of them is stationary, then an error-correction representation of the variables is allowed (Engle and Granger (1987), Johansen (1988)). The correspondence between cointegrating vectors and error-correction models is also the basis for the two-step estimation procedure proposed by Engle and Granger. Instead of a simultaneous specification of the long-run and short-run properties of the variables, the two-step procedure sequentially models the long-run equilibrium relation and the short-run dynamics.

In the first step, after pre-testing the variables entering the cointegrating relation in order to ensure that they are of the same order of integration, an estimate of the cointegrating vector is obtained by means of a static OLS regression (Engle and Granger

[3] An additional rationale for this result is provided by Nickell (1985) in terms of aggregation problems. If there are two groups of agents with identical targets but different adjustment cost parameters (or if identical agents are adjusting two components of the choice variable to the same target but with different costs), it can be shown that the optimal decision rule involves, after aggregation, an error correction term with a negative coefficient.

(1987)) or by application of Johansen's (1988, 1991) maximum likelihood method if more than two variables are involved. In the second step, the residuals from the estimated cointegrating relations are used as an error-correction term and a *general to specific* specification strategy can be applied to model the dynamics of all variables around the already determined equilibrium relation.

In the empirical analysis of the next section the focus will be on the simultaneous specification of the long- and short-run features of the data, though also the two-step procedure will be applied and the results compared.

2.2. Modelling the Italian demand for M2.

The general to specific econometric methodology outlined above is applied here to the specification of a money demand function for Italy over the period 1964-1986. The series included in the analysis are quarterly, seasonally unadjusted from 1962(1) to 1986(2) and are defined as follows (interest rates are expressed as fractions):[4]

m : (log of) end-of-period stock of M2 held by the public. M2 includes notes and coins, bank and postal current and deposit accounts, and interest bearing postal bills (*Buoni Fruttiferi Postali*);

p : (log of) GDP deflator;

y : (log of) 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 a weighted average of government bonds and private bonds before 1974 and of government bonds, private bonds and Treasury Bills (*Buoni Ordinari del Tesoro, BOT*) from 1974 onwards. The weights are determined by end-of-period outstanding stocks;

S_i: seasonal dummies;

D83q4: dummy variable (taking the value of 1 in 1983(4) and 0 everywhere else), introduced to eliminate the effect of a statistical anomaly in the data for M2 due to lags in data collection on the amount of bank deposits in December 1983.

The underlying theoretical model of the demand for money is in accord with standard theory, with a scale variable (real income) and a set of relevant yields on money

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[4] Data sources are given in the *Appendix*. The sample period ends in 1986(2), the last available observation for the series of GDP used in the analysis. After 1986, a new series for GDP has been published by the Italian Central Statistical Office (ISTAT). However, no reconstruction of the quarterly series is available for the 1960s and 1970s; therefore, we decided to use the old GDP series, which is homogeneous throughout the whole sample period.

and alternative assets as determinants of real money holdings.

Starting with the simultaneous specification of the long-run equilibrium and the short-run dynamics, we estimate a general baseline model with five lags of all variables, except the price level, nine lags of which are included to allow for a potential fifth-lag effect of the annual rate of inflation. The baseline model is therefore the following:

$$\begin{aligned}(m-p)_t = c + \sum_{i=1}^{5} \alpha_i (m-p)_{t-i} + \sum_{i=0}^{5} \beta_i y_{t-i} + \sum_{i=0}^{5} \gamma_{1i} R^b_{t-i} + \sum_{i=0}^{5} \gamma_{2i} R^m_{t-i} + \sum_{i=0}^{9} \mu_i p_{t-i} + \\ + \theta_1 S_{1t} + \theta_2 S_{2t} + \theta_3 S_{3t} + \theta_4 D83q4_t + \varepsilon_t\end{aligned} \tag{11}$$

The results from estimation of this baseline equation are reported in Table 1, together with a set of diagnostic tests designed to evaluate the congruency of the model with the data. Residual serial correlation up to the fifth order and normality are tested by means of the Lagrange Multiplier test *AR(5)* and the Jarque-Bera statistic respectively. Engle's (1982) *ARCH(4)* test for fourth-order autoregressive conditional heteroscedasticity and the *RESET* test for the correct specification of the linear form against a quadratic alternative are also reported.

The estimation of the baseline model delivers a standard error of the regression of 0.94%, which is acceptably low if compared with other money demand studies conducted on a similar sample period (for example, Vaciago and Verga (1989) report residual standard errors of about 1.1-1.2%[5]) and the diagnostic tests do not detect any sign of mis-specification. However, in the light of the work by Baba *et al.*(1992) on U.S. data, suggesting that careful modelling of interest rate volatility may substantially improve the performance of estimated money demand equations, proxies for volatility were tried as additional regressors in the baseline model. The chosen proxies were the four-quarter moving standard deviations (*MSD*) of the two interest rates R^b and R^m, calculated as:

[5] In another recent study, Muscatelli and Papi (1990) focused on the modelling of the learning process of wealth holders when new financial instruments are introduced. The standard error of their preferred equation, capturing the learning process by means of a logistic-type trend included in the long-run money demand equation, is remarkably low (0.3%). In our analysis, some part of the financial innovation effect may be captured by the measure of the alternative interest rate R^b, which includes the yield on Treasury Bills (*BOT*) - whose introduction represents the main financial innovation of the 1970s - with a growing weight as the proportion of total portfolios invested in them increases. Unfortunately, Muscatelli and Papi do not report structural stability results from recursive estimation of their model, which makes their investigation, though interesting, not directly comparable with ours.

Table 1

Feedback specification : baseline model.

Dependent variable: $(m\text{-}p)_t$ Sample period: 1964(2)-1986(2)
(standard errors in parentheses)

Variable	lag *i* 0	1	2	3	4	5
$(m\text{-}p)_{t\text{-}i}$	-	0.638	0.189	0.086	0.746	-0.630
		(0.128)	(0.141)	(0.137)	(0.153)	(0.126)
$y_{t\text{-}i}$	0.299	-0.118	-0.190	0.082	-0.219	0.156
	(0.097)	(0.110)	(0.100)	(0.104)	(0.110)	(0.092)
$R^b_{t\text{-}i}$	0.169	-0.664	0.013	0.431	0.253	0.428
	(0.212)	(0.332)	(0.392)	(0.402)	(0.396)	(0.320)
$R^m_{t\text{-}i}$	-0.220	0.773	0.109	-0.237	-0.748	-0.322
	(0.441)	(0.534)	(0.516)	(0.514)	(0.521)	(0.421)
$p_{t\text{-}i}$	-0.991	0.557	0.163	0.398	0.678	-0.720
	(0.102)	(0.191)	(0.204)	(0.196)	(0.203)	(0.168)
$p_{t\text{-}6\text{-}i}$	0.041	0.003	-0.074	-0.076		
	(0.113)	(0.112)	(0.119)	(0.093)		
constant	-0.072					
	(0.142)					
$D1983q4_t$	-0.047					
	(0.014)					
S_{it}		-0.011	-0.031	-0.055		
		(0.025)	(0.024)	(0.022)		

$R^2 = 0.9994$ $\sigma = 0.942\%$

Diagnostic tests [p-value]:

AR(5) :	$F(5,46) = 1.11$ [0.37]	*Normality* :	$\chi^2(2) = 3.96$ [0.14]
ARCH(4) :	$F(4,43) = 0.23$ [0.92]	*RESET* :	$F(1,50) = 3.60$ [0.07]

Note: S_i ($i=1,2,3$) denote quarterly dummy variables. σ is the estimated standard error of the regression. *AR*(5) is the *F*-version of the Lagrange multiplier test for residual serial correlation up to the 5th order; *Normality* χ^2 is the Jarque-Bera test for residual normality; *ARCH*(4) is the test for autoregressive conditional heteroscedasticity up to the 4th order in *F*-form (Engle (1982)); *RESET* is the *F*-version of the regression specification test (functional form).

$$MSD(R^k)_t = \sqrt{\frac{\sum_{j=0}^{4}\left(R^k_{t-j} - MA(R^k)_t\right)^2}{5}} \quad with \quad MA(R^k)_t = \frac{\sum_{i=0}^{4} R^k_{t-i}}{5} \quad and \quad k = b,m.$$

The extended baseline model showed no improvement in the residual standard error and the coefficients on current and lagged *MSD* terms were not statistically significant. Therefore, we retained the original model in Table 1 as the starting point for a process of reduction, reparameterization and testing in order to reach a more parsimonious representation of the Data Generating Process. An intermediate stage in the specification process shows that dynamics only of the first, fourth and fifth order (the latter only for the *m-p* and *y* series) are relevant and the annual inflation rate ($\Delta_4 p_t$) is the only variable involving prices which enters the equation with some explanatory power at time *t* and *t-1*. 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 by the data. Difference restrictions of the appropriate order are then imposed on all variables, leaving the levels (lagged one period) of *m-p*, *y*, R^b, R^m and $\Delta_4 p$ to capture the long-run solution of the equation. The estimated long-run relation (commented below) between money balances and its determinants is the following (coefficient standard errors are reported in parentheses):

$$\underset{}{m - p} = \underset{(0.163)}{1.578\, y} - \underset{(1.990)}{1.469\, R^b} + \underset{(3.850)}{2.531\, R^m} + \underset{(0.497)}{2.062\, \Delta_4 p} \qquad (12)$$

Residuals from the above long-run solution form the *ECM* term which, lagged one period, enters the final specification of the feedback model:

$$\begin{aligned}\Delta_4(m-p)_t = {} & \underset{(0.060)}{0.730\, \Delta_4(m-p)_{t-1}} + \underset{(0.064)}{0.171\, \Delta_4 y} - \underset{(0.059)}{0.201\, \Delta_4 y_{t-1}} - \underset{(0.068)}{0.951\, \Delta\Delta_4 p_t} \\ & - \underset{(0.123)}{0.784\, \Delta_3 R^b_{t-1}} + \underset{(0.197)}{0.730\, \Delta_3 R^m_{t-1}} + \underset{(0.016)}{0.083\, ECM_{t-1}} \\ & - \underset{(0.010)}{0.056\, D1983q4_t} - \underset{(0.003)}{0.003\, S_{1t}} - \underset{(0.004)}{0.018\, S_{2t}} - \underset{(0.003)}{0.006\, S_{3t}} + \underset{(0.036)}{0.189}\end{aligned} \qquad (13)$$

$R^2 = 0.9975$ $\sigma = 0.953\%$

Diagnostic tests [p-value]:

Durbin's h =	- 0.94 [0.35]	*AR*(5) :	*F*(5,72) = 0.53 [0.75]
Normality :	$\chi^2(2)$ = 5.90 [0.06]	*ARCH*(4) :	*F*(4,69) = 0.32 [0.86]
Heterosc. :	*F*(19,57) = 0.81 [0.68]	*RESET* :	*F*(1,76) = 1.36 [0.25]

In addition to the diagnostic tests performed on the baseline model, residual unconditional heteroscedasticity due to the squares of the regressors is tested (White (1980)). When tested against the unrestricted baseline model, the 26 parameter restrictions embodied in the final specification are not rejected: the $F(26,51)$ statistic is 1.06, with a p-value of 0.42.[6] The standard error of the regression is satisfactorily low (0.95%) and there is no sign of residual autocorrelation, heteroscedasticity and functional form mis-specification. Only the normality test yields a value of 5.90, close to the 5% critical value. Also the forecasting performance of the model has been evaluated by respecifying the equation over the sample up to 1984(2), with the last eight observations left for the forecasting analysis. The same variables as in equation (13) entered the final specification on the shorter sample and the value of the Chow test for parameter constancy over the 1984(3)-1986(2) period was $F(8,67)=0.75$ (5% critical value 2.10).

Several features of the final equation deserve some comment, prior to assessing its structural stability performance. The dependent variable is expressed as a fourth-order difference, but qualitatively very similar results in terms of the long-run solutions and the overall performance of the equation are obtained when the first-order difference of real money balances is used as the dependent variable (the *Appendix* reports the final estimated equation in this case). The resulting short-run dynamics involves differences of the regressors of several orders, with a sizeable negative effect of the acceleration of (annual) inflation. Lagged third-order differences of both interest rates enter the equation with coefficients of opposite sign and similar magnitudes, possibly capturing the effect of the changing interest rate variability over the sample period.[7] The long-run solution in (12) displays an elasticity of real money holdings to income well above unity (1.58), a positive semi-elasticity to the inflation rate, and different (but correctly signed) semi-elasticities to R^b and R^m, with the latter larger in absolute value. Although the high standard errors do not allow sharp inferences on the values of the long-run interest rate elasticities (indeed, the hypothesis of a long-run money demand independent of interest rates cannot be rejected), the different long-run responses of m-p to R^b and R^m, together with the positive sign of the elasticity to the inflation rate, are two features of the final specification which seem difficult to justify on standard theoretical grounds and may suggest an explanation based on the non-structural nature of the estimated long-run parameters when expectations formation is a relevant, but neglected, aspect of agents' behaviour, as discussed in section 3.3 of the

[6] When the test is performed keeping the unconstrained lagged levels of the variables in the regression (before imposing the *ECM* term) the resulting $F(22,51)$ statistic is 1.25 (0.25).

[7] As a further check on the irrelevance of more specifically designed measures of interest rate volatility, we added to the final specification the contemporaneous and lagged values of the $MSD(R^b)$ and $MSD(R^m)$ variables defined previously; again, no significant volatility effect was detected.

preceding chapter (Kelly (1985), Cuthbertson (1986a)).[8] An additional feature of equation (13) which might suggest an alternative underlying model based on expectations is the positive coefficient on the lagged long-run relation, implying an amplification of past deviations from the equilibrium instead of an error-correcting behaviour. Similar results are not uncommon in applied studies on Italian money demand. Muscatelli (1991), using both single-equation and system estimation methods over the 1966-1987 period, reports a positive coefficient on the lagged level term in the equation for real money balances. However, estimation of a complete system of equations for money, income, interest rates and inflation shows that the reactions of the variables other than money to deviations from the long-run equilibrium ensure the dynamic stability of the system as a whole, notwithstanding the apparent instability of the dynamic adjustment of money balances. A feedback, error-correction interpretation of the short-run dynamics may then be validly applied to the complete system, with no appeal to a forward-looking alternative structure.

The property of balance of the final specification is checked by testing for the order of integration of the variables included in equation (13). The results, reported in the *Appendix*, show that the hypothesis of non-stationarity is clearly rejected for all variables except the *ECM* term, for which the Dickey-Fuller test does reject non-stationarity only at the 10% level.[9]

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[8] To illustrate such theoretical possibility, consider the following very simplified representation of a feedback model for money demand:

$$(m-p)_t = b_1 y_t - b_2 R^b_{t-1} + b_3 R^m_{t-1} + b_4 \pi_{t-1} + \varepsilon_t \quad , \quad \textit{with } 0 < b_2 < b_3 \quad \textit{and } b_4 > 0$$

where the pattern of coefficients is consistent with a positive long-run elasticity to inflation and a semi-elasticity to R^m higher than that with respect to R^b. Now suppose that the underlying model is a forward-looking one, specified as follows:

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

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 processes for R^b and R^m can be represented as:

$$R^{be}_t = \beta_1 R^b_{t-1} + \beta_2 \pi_{t-1} \quad (\beta_i > 0) \quad , \quad R^{me}_t = \delta_1 R^m_{t-1} + \delta_2 \pi_{t-1} \quad (\delta_i > 0)$$

then the feedback specification can be interpreted as the reduced form of the forward-looking money demand and the expectations generating equations and the following restrictions would hold: $b_2 = \alpha\beta_1$, $b_3 = \alpha\delta_1$ and $b_4 = \alpha(\delta_2 - \beta_2)$. Therefore, estimates of the long-run elasticities that appear difficult to justify on theoretical grounds could well be generated from the above model if $\delta_1 > \beta_1$ and $\delta_2 > \beta_2$.

[9] However, Kremers, Ericsson and Dolado (1992) show that results from *DF* (and augmented *DF*) tests must be interpreted with care, since the testing procedure imposes a common factor restriction in the regression used to implement the test and the inability of rejecting non-stationarity may be due to dynamic mis-specification. In principle, this problem could be overcome by using

Finally, the structural stability of the final specification is tested by means of recursive stability tests. The main results are summarized in Figure 1. Panel (*a*) plots the one-step innovations, defined as $v_t=y_t-x_t'\beta_{t-1}$ (where $y_t=\Delta_4(m-p)_t$, x_t is the vector of regressors at time t and β_{t-1} is the vector of coefficients estimated over the sample ending at time $t-1$), and panel (*b*) shows the recursive residuals $u_t=y_t-x_t'\beta_t$ with two standard error bands. Several large innovations are detected in the first half of the 1970s, starting in 1970. Correspondingly, the recursive standard error of the regression, used to construct the confidence interval for the residuals u_t, increases sharply at the beginning of 1970 and again in 1973/74. These results suggest instability of the estimated coefficients and of the regression error variance at various dates in the first part of the 1970s, formally evaluated by means of recursive one-step and break-point Chow (1960) stability tests (*Chow1* and *ChowN*) calculated at each date from T_1+1, with T_1 being the last observation used for initialization of the recursive procedure, to the end of the sample T. Formally, being RSS_t the residual sum of squares up to time t and k the number of regressors in the equation, we have:

$$Chow1_t = \frac{RSS_t - RSS_{t-1}}{\frac{RSS_{t-1}}{(t-1)-k}} = \frac{v_t^2}{\hat{\sigma}_{t-1}^2\,\omega_t^2} \sim F(1,(t-1)-k)$$

$$ChowN_t = \frac{\frac{RSS_T - RSS_{t-1}}{T-(t-1)}}{\frac{RSS_{t-1}}{(t-1)-k}} = \frac{\sum_{i=t}^{T}\frac{v_i^2}{\hat{\sigma}_{t-1}^2\,\omega_i^2}}{T-(t-1)} \sim F(T-(t-1),(t-1)-k)$$

where $\omega^2_t=1+x_t'(X_{t-1}'X_{t-1})^{-1}x_t$ (with X_{t-1} being the regressor matrix up to $t-1$), so that $\sigma^2_{t-1}\omega^2_t$ is the variance of the one-step forecast error between $t-1$ and t. The outcome of the tests is plotted in panels (*c*) and (*d*), scaled by the 5% critical values from the appropriate F distributions. Instability in the first half of the 1970s is confirmed and the structural break in 1970 stands out as the major episode of this kind. Finally, panels (*e*)-(*h*) assess the stability of the recursively estimated coefficients on the regressors in (13) involving income, inflation and the interest rates. All coefficients display remarkable stability from the mid-'70s onwards, with high variability concentrated in the first half of the 1970s. In particular, the coefficients on Δ_3R^b and Δ_3R^m show sharp changes at the beginning of 1970. This finding suggests that the specification in equation 3, though acceptable on the basis of standard diagnostic tests, fails to capture some important features of money demand behaviour.

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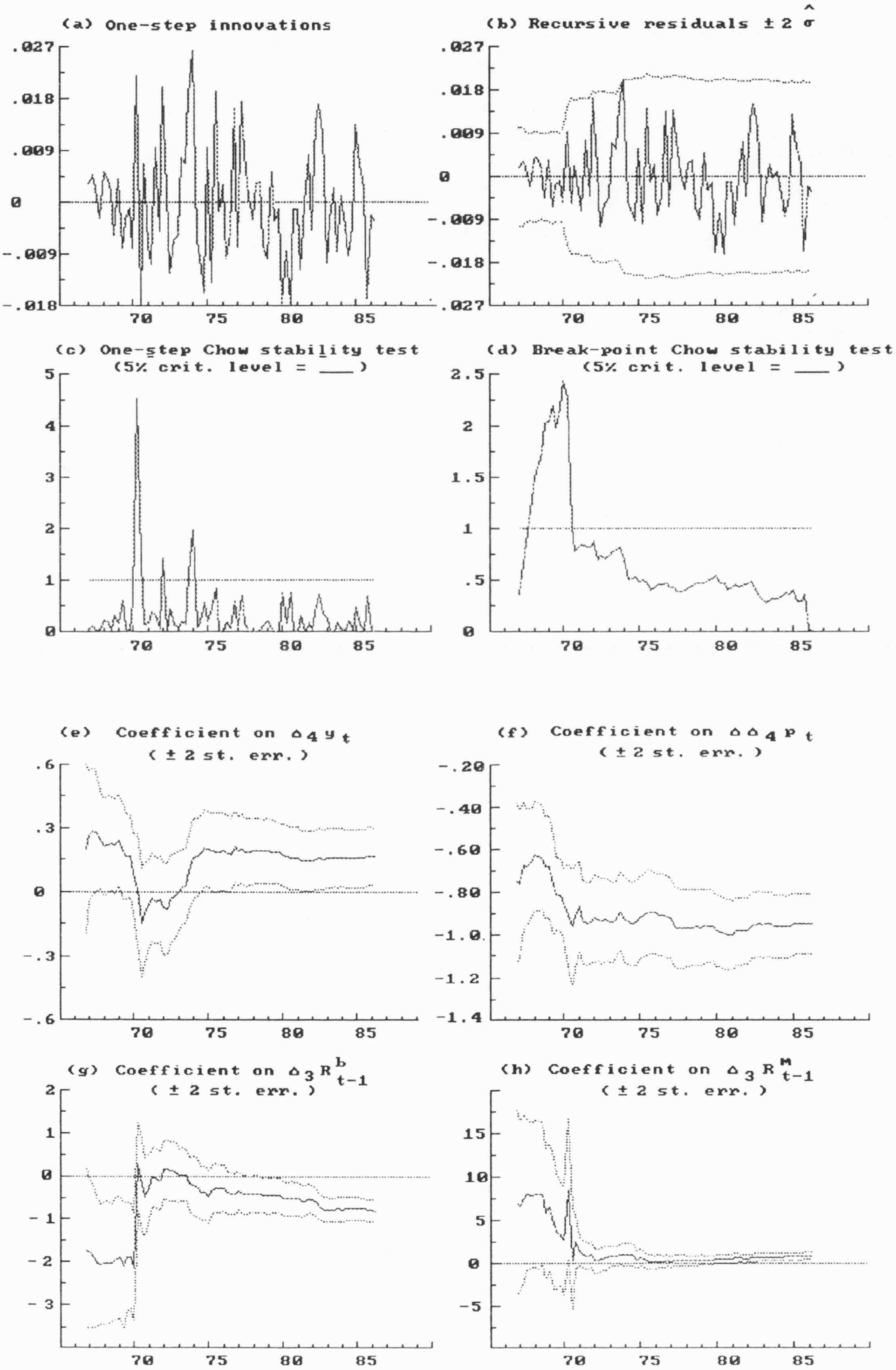
directly the t-statistic on the *ECM* term in the final feedback equation modelling an *I(0)* variable to evaluate non-stationarity, since such specification does not impose any common factor restriction.

Figure 1
Stability analysis of the feedback model



Before trying to relate such episodes of instability to the processes generating the determinants of money demand, we investigated whether similar results may be obtained from a feedback model specified using cointegration methods. The *Appendix* reports the outcome of the implementation of the two-step Engle-Granger specification procedure (Engle and Granger (1987)) and of the Johansen (1988, 1991) maximum likelihood method for estimating cointegrating vectors. Here we briefly summarize the main results.

The Dickey-Fuller test performed on the residuals from the static OLS cointegrating regression (with real money balances as dependent variable) suggests cointegration among the series. The estimated long-run elasticities to income and inflation are similar to those previously found in (12), whereas now the semi-elasticity to R^b is larger, in absolute value, than that to R^m. When included in the dynamic specification of the feedback model, the lagged residuals from the cointegrating regression have a zero coefficient. The other features of equation (13) are qualitatively confirmed, including the instability pattern (the only difference being the greater importance of the break around 1973/74 relative to the episode in 1970). The Johansen's procedure, implemented on a fifth-order, five-variable vector autoregression, yields evidence of two valid cointegrating vectors in the system; the estimated coefficients allow a possible money demand interpretation only for the first vector (though with the already detected positive inflation effect), whereas the second displays "wrongly" signed interest rate responses. In view of the temporal instability of the underlying *VAR*, the results from the Johansen's procedure must be considered with extreme care, since the estimates of both the cointegration rank and the coefficients of the cointegrating vectors may be unstable over the sample.[10]

This concludes the empirical analysis of a feedback money demand equation for Italy. In the next section our main results concerning the structural instability of the final specification presented above will be reconsidered and related to the time-series behaviour of the determinants of money holdings.

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[10] In the next chapter the multiplicity of cointegrating vectors will be the focus of the analysis; in that case, however, the stability of the underlying *VAR* system is verified over the relevant sample period (1983-1991) and more reliable inferences can be drawn from the estimates.

3. *Structural instability and the behaviour of the determinants of money demand.*

Two main results from the empirical analysis of the preceding section suggest that the feedback money demand equations presented may not adequately characterize underlying agents' behaviour. First, the coefficient on the error-correction term in the final specification, which should capture agents' reaction to past deviations from the long-run equilibrium, is positive, not supporting a feedback, rule-of-thumb interpretation. Second, the structural stability analysis reveals that the feedback equation, however specified, suffers from remarkable instability at various dates in the sample. As argued in the previous chapter, structural instability of feedback specifications may be caused by the neglect of one important aspect of agents' behaviour, namely expectations formation. If individuals choose their money holdings on the basis of (rational) expectations concerning the future evolution of real income, interest rates and inflation, the instability of the estimated feedback equation may be due to shifts in the expectations generating processes. Specific causes for such shifts may be perceived changes in policy regimes or sharp alterations in the time-series behaviour of some relevant variables.

This possibility is investigated in the present section using formal tests for superexogeneity and invariance (Engle and Hendry (1993)). Several applications of this testing procedure may be found in the recent money demand literature. Fischer and Peytrignet (1991) refuted the practical relevance of the Lucas (1976) critique for a feedback specification of money demand for Switzerland, studying its stability in the face of repeated monetary policy regime changes (from M1 targeting to exchange rate management and then to monetary base targeting) in the 1970s and 1980s. Qualitatively similar results are reported by Hurn and Muscatelli (1992) for the demand for a broad U.K. money aggregate (M4) and by Hendry and Ericsson (1991) and Hendry and Engle (1993) for the U.K. narrow money (M1) demand function.[11]

3.1. The testing framework.

To evaluate formally the dependence of the instability in the feedback model onto changes in the time-series behaviour of the regressors we adopt the framework of Engle and Hendry (1993) and Hendry (1994) to test for superexogeneity and invariance in conditional models.

Given the joint density of generic variables y_t and x_t, $D(y_t,x_t|I_t\ ;\lambda_t)$, where I_t is an

[11] In a different context, Fischer (1989) applies similar testing procedures to assess the invariance of monetary expectations to the policy regime shifts occurred in the U.S. in the 1979/82 period.

information set including valid (current and past) conditioning variables and λ_t is a vector of parameters, the following factorization is always possible:

$$D(y_t, x_t \mid I_t; \lambda_t) = D_{y|x}(y_t \mid x_t, I_t; \lambda_{1t}) \cdot D_x(x_t \mid I_t; \lambda_{2t}) \tag{14}$$

where $D_{y|x}$ and D_x are the conditional density of y_t given x_t and the marginal density of x_t respectively, and λ_{1t}, λ_{2t} are the corresponding parameter vectors. Adopting a special, but operationally useful, case let us assume that y and x are jointly normally distributed with (possibly time-dependent) mean vector μ_t and variance matrix Σ_t, so that $\lambda_t = (\mu_t^y, \mu_t^x, \sigma_t^{yy}, \sigma_t^{xx}, \sigma_t^{yx})'$ yielding the following conditional relation:

$$\begin{aligned} & y_t \mid x_t, I_t \sim N\left[\delta_t(x_t - \mu_t^x) + \mu_t^y, \omega_t\right] \\ & \textit{with} \quad \delta_t = \frac{\sigma_t^{yx}}{\sigma_t^{xx}} \quad , \quad \omega_t = \sigma_t^{yy} - \frac{(\sigma_t^{yx})^2}{\sigma_t^{xx}} \end{aligned} \tag{15}$$

Ignoring other regressors, the behavioural relation to be modelled is $\mu_t^y = \beta_t(\lambda_{2t}) \cdot \mu_t^x$, where β is allowed to vary over time and in response to changes in the parameters of the marginal density of x_t, $\lambda_{2t} = (\mu_t^x, \sigma_t^{xx})'$. We can now state the conditions under which one may estimate a valid regression model of the form: $y_t = \beta x_t + \epsilon_t$ where $\epsilon_t \sim N(0, \omega)$. Such conditions concern the weak exogeneity of x for the parameters of interest, constancy of the regression coefficients and invariance of β to changes in the elements of λ_{2t}. Using the theoretical relation between μ_t^y and μ_t^x, the conditional model (15) may be expressed as:

$$y_t \mid x_t, I_t \sim N\left[\beta_t(\mu_t^x, \sigma_t^{xx}) \cdot x_t + \left(\delta_t - \beta_t(\cdot)\right) \cdot (x_t - \mu_t^x) \ , \ \omega_t\right] \tag{16}$$

Weak exogeneity of x_t for β requires that the parameters of the marginal model (μ_t^x and σ_t^{xx}) do not enter the conditional model, so that there is no loss of information about β_t from not modelling the marginal model for x_t. Necessary condition is $\delta_t = \beta_t(\lambda_{2t})$. *Constancy* of the regression coefficients requires $\delta_t = \delta \ \forall t$; moreover, given the definition of δ_t in (15), it must be that $\sigma_t^{yy} = \omega + \delta \sigma_t^{yx}$ in order to have a homoscedastic conditional model with variance ω. Finally, *invariance* of β to changes of λ_{2t} obtains when $\beta_t(\lambda_{2t}) = \beta_t \ \forall t$, so that possible parameter variations over time do not depend on modifications of μ_t^x and σ_t^{xx}. If weak exogeneity and invariance jointly hold, then x_t is *superexogenous* for the parameter of interest β.

To implement a test for invariance and superexogeneity, the alternative hypothesis of changes in λ_2 determining variations in β must be made explicit. Engle and Hendry (1993) adopt the following approximation (assuming $\mu_t^x \neq 0$):

$$\beta_t(\mu_t^x, \sigma_t^{xx}) = \beta_0 + \beta_1 \mu_t^x + \beta_2 \frac{\sigma_t^{xx}}{\mu_t^x} + \beta_3 \sigma_t^{xx} \tag{17}$$

where the moments of the marginal distribution of x influence β with a time-invariant relationship, since the β_i's are assumed independent of time; invariance entails $\beta_1 = \beta_2 = \beta_3 = 0$. Using (17) and the expansion $\delta_t = (\sigma_t^{yx}/\sigma_t^{xx}) = \delta_0 + \delta_1 \sigma_t^{xx}$ to allow for potential non-constancy of the regression coefficient, the conditional model (16) may be rewritten as:

$$y_t \mid x_t, I_t \sim N\left[\beta_0 x_t + (\delta_0 + \delta_1 \sigma_t^{xx} - \beta_0) \cdot (x_t - \mu_t^x) + \beta_1 (\mu_t^x)^2 + \beta_2 \sigma_t^{xx} + \beta_3 \sigma_t^{xx} \mu_t^x \ , \ \omega_t\right] \tag{18}$$

To reach a testable form for (18), μ_t^x and σ_t^{xx} must be parameterized from estimation of a reduced form model for x_t, $x_t = \pi_x' z_t + \eta_t$, where z_t is a set of valid instruments for x_t and allows for regime shifts and other sources of structural change (e.g. by means of dummy variables). Fitted values and residuals from this model may then be used as measures of μ_t^x and $(x_t - \mu_t^x)$ respectively; functions of the estimated residuals, such as a moving average of squared residuals, may be employed to construct a series for σ_t^{xx}. The tests can then be performed on the following regression model:

$$y_t = \beta_0 x_t + (\delta_0 - \beta_0)\hat{\eta}_t + \delta_1(\hat{\sigma}_t^{xx} \cdot \hat{\eta}_t) + \beta_1 \hat{x}_t^2 + \beta_2 \hat{\sigma}_t^{xx} + \beta_3(\hat{\sigma}_t^{xx} \cdot \hat{x}_t) + \varepsilon_t \tag{19}$$

Now a zero estimated coefficient on $\hat{\eta}_t$ entails weak exogeneity of x_t for β and a zero coefficient on $(\hat{\sigma}_t^{xx}\hat{\eta}_t)$ imply constancy (corresponding to $\delta_1 = 0$ in the expansion of δ_t). Invariance, implying $\beta_1 = \beta_2 = \beta_3 = 0$, is tested on the coefficients of the remaining regressors. Moreover, under superexogeneity, the determinants of regime shifts in the marginal model for x_t do not have any influence on the conditional model; therefore a direct test of superexogeneity of x_t can be conducted by adding those variables, included in z_t above, capturing structural change in x_t, to the conditional model and verifying that they are statistically insignificant.

In what follows the general testing framework outlined above is applied to our feedback model for money demand. Three steps are involved. First, we assess the stability of reduced form models for the regressors in the feedback specification (equation (13)), informally relating the results to the instability pattern found for the feedback model in the previous section. The estimated models are then extended with the inclusion of additional variables capturing structural changes in order to attain a reasonably stable formulation for the marginal models; at this stage empirical measures for the moments of the regressors distributions are constructed. Finally, invariance and superexogeneity tests are performed on the feedback specification to formally evaluate the dependence of the detected instability on shifts in the process generating the regressors.

3.2. An application to Italian money demand.

The specification of a marginal model for the regressors in (13) ($\Delta_4 y$, $\Delta\Delta_4 p$, $\Delta_3 R^b$ and $\Delta_3 R^m$) begins from a general baseline equation for each variable, which is then simplified and reparameterized so as to reach a parsimonious representation of the variable of interest. As for the feedback model of the previous section, a set of diagnostic tests is used to assess the adequacy of the final specifications, and the imposed parameter restrictions are tested against the initial baseline models. Finally, a stability analysis is performed by means of recursive tests. Since, as noted by Cuthbertson (1991) (see also chapter 1, section 3.3), it may easily be the case that, using a limited set of variables in a finite sample, instability is detected even in the absence of true structural shifts in the underlying process generating the dependent variable, we concentrate on the apparently most relevant instability episodes. For this reason we employ a 1% critical level in the implementation of the recursive stability tests of this section. Table 2 reports the final specifications together with diagnostic tests and general tests of parameter restrictions. Figures 3 to 6 show the results of the stability analysis on the estimated models.

The baseline model for the annual rate of growth of GDP, $\Delta_4 y$, includes five lags of the dependent variable, of the annual rate of change of real money balances, of inflation and of the levels of the two interest rates R^b and R^m. Also a linear time trend and seasonal dummies enter the equation. The final specification allows for an overall negative effect of past inflation changes and a positive effect of past real money growth (Table 2, equation 1); no instability is detected (Figure 3). Modelling the acceleration of the inflation rate, $\Delta\Delta_4 p$, proved more difficult, given the time-series behaviour of the annual rate of inflation (Figure 2, panel (*a*)), showing several local peaks in the 1970s and early 1980s, reflected in large and repeated swings in its first difference (panel (*b*)). Obtained as a reduction of a baseline model with four lags of the dependent variable and five of the rate of real money balances and output and of the interest rate levels, the final equation features a sizeable effect of past acceleration in $\Delta_4(m-p)$ (equation 2). Although the break-point Chow test does not indicate any major episode of instability, the recursive residuals and the one-step stability test detect two serious breaks in 1969 and again in 1973, with some minor sign of instability also in 1971 (Figure 4). The time-series behaviour of interest rates is plotted in Figure 2, panel (*c*): in several periods the level of the alternative rate, R^b, sharply rises and then rapidly declines, and the own-yield on money, R^m, follows a similar pattern but with less pronounced fluctuations. These characteristics are reflected in the behaviour of the interest rate

Table 2
Marginal models for $\Delta_4 y$, $\Delta\Delta_4 p$, $\Delta_3 R^b$, $\Delta_3 R^m$

1. $\Delta_4 y_t$

$$\Delta_4 y_t = \underset{(0.071)}{0.623}\,\Delta_4 y_{t-1} - \underset{(0.058)}{0.285}\,\Delta_4 y_{t-4} + \underset{(0.038)}{0.229}\,\Delta_4(m-p)_{t-1} + \underset{(0.146)}{0.326}\,\Delta_3 R^b_{t-1}$$

$$- \underset{(0.246)}{0.569}\,\Delta_4 R^m_{t-1} - \underset{(0.159)}{0.476}\,\Delta\Delta_4 p_{t-3} + \underset{(0.076)}{0.190}\,\Delta_2\Delta_4 p_{t-1} + \underset{(0.003)}{0.012}$$

	$R^2 = 0.827$	$\sigma = 1.398\%$	
$AR(5)$:	$F(5,75) = 0.81$ [0.55]	*Normality* :	$\chi^2(2) = 0.14$ [0.93]
$ARCH(4)$:	$F(4,72) = 2.12$ [0.09]	*Heterosc.*:	$F(16,63) = 1.13$ [0.35]
$X_i{*}X_j$:	$F(44,35) = 1.10$ [0.39]	*RESET* :	$F(1,79) = 0.08$ [0.78]

Test of restrictions against "baseline" model: $F(21,59) = 0.55$ [0.93]

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2. $\Delta\Delta_4 p_t$

$$\Delta\Delta_4 p_t = \underset{(0.039)}{0.084}\,\Delta_4(m-p)_{t-1} + \underset{(0.072)}{0.332}\,\Delta_4(m-p)_{t-4} - \underset{(0.063)}{0.334}\,\Delta_4(m-p)_{t-5} + \underset{(0.166)}{0.384}\,\Delta_2 R^b_{t-1}$$

$$+ \underset{(0.249)}{0.410}\,\Delta_3 R^m_{t-1} - \underset{(0.002)}{0.005}$$

	$R^2 = 0.417$	$\sigma = 1.478\%$	
$AR(5)$:	$F(5,78) = 1.35$ [0.25]	*Normality* :	$\chi^2(2) = 2.16$ [0.34]
$ARCH(4)$:	$F(4,75) = 0.36$ [0.84]	*Heterosc.*:	$F(10,72) = 0.92$ [0.52]
$X_i{*}X_j$:	$F(20,62) = 0.80$ [0.70]	*RESET* :	$F(1,82) = 0.23$ [0.63]

Test of restrictions against "baseline" model: $F(22,61) = 1.05$ [0.43]

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3. $\Delta_3 R^b_t$

$$\Delta_3 R^b_t = \underset{(0.180)}{0.374}\,\Delta R^b_{t-1} + \underset{(0.111)}{0.975}\,\Delta_2 R^b_{t-1} - \underset{(0.034)}{0.121}\,R^b_{t-1} - \underset{(0.029)}{0.066}\,\Delta_4(m-p)_{t-3}$$

$$+ \underset{(0.021)}{0.063}\,\Delta(m-p)_{t-4} + \underset{(0.036)}{0.090}\,\Delta_4 y_{t-1} + \underset{(0.025)}{0.074}\,\Delta_4 p_{t-4} + \underset{(0.035)}{0.047}\,\Delta_2\Delta_4 p_{t-3} + \underset{(0.436)}{0.536}$$

	$R^2 = 0.852$	$\sigma = 0.663$	
$AR(5)$:	$F(5,75) = 1.19$ [0.32]	*Normality* :	$\chi^2(2) = 5.64$ [0.06]
$ARCH(4)$:	$F(4,72) = 0.70$ [0.59]	*Heterosc.*:	$F(16,63) = 1.14$ [0.34]
$X_i{*}X_j$:	$F(44,35) = 1.46$ [0.13]	*RESET* :	$F(1,79) = 1.31$ [0.26]

Test of restrictions against "baseline" model: $F(25,55) = 1.24$ [0.24]

Table 2/contd.

4. $\Delta_3 R^m_t$

$$\Delta_3 R^m_t = \underset{(0.087)}{0.841}\ \Delta_2 R^m_{t-1} - \underset{(0.019)}{0.047}\ R^m_{t-1} + \underset{(0.070)}{0.272}\ \Delta R^b_{t-1} + \underset{(0.046)}{0.097}\ \Delta_3 R^b_{t-1}$$

$$- \underset{(0.010)}{0.032}\ \Delta_4 (m-p)_{t-5} + \underset{(0.015)}{0.045}\ \Delta_4 y_{t-1} + \underset{(0.153)}{0.303}$$

	$R^2 = 0.885$	$\sigma = 0.320$	
AR(5) :	$F(5,77) = 0.20$ [0.96]	*Normality* :	$\chi^2(2) = 3.43$ [0.18]
ARCH(4) :	$F(4,74) = 2.60$ [0.04]	*Heterosc.*:	$F(12,69) = 2.23$ [0.02]
$X_i{*}X_j$:	$F(27,54) = 1.81$ [0.03]	*RESET* :	$F(1,81) = 6.12$ [0.02]

Test of restrictions against "baseline" model: $F(27,55) = 1.39$ [0.15]

Note: The sample period is 1964(2)-1986(2). $X_i{*}X_j$ is a general test for heteroscedasticity related to the squares and cross-products of the regressors; the remaining diagnostic tests are illustrated in the notes to table 1. For each variable, the "baseline" model represents the unrestricted dynamic model from which the specification search started.

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regressors in (13), $\Delta_3 R^b$ and $\Delta_3 R^m$, with several large values of either sign at the beginning of the 1970s and in the middle of the decade, and again in the early 1980s (panel (*d*)). Modelling of the time-series behaviour of the third-order differences of R^b and R^m starts with identical baseline equations, including five lags of the interest rates and inflation, and nine lagged levels of real money balances and income. Differences of various orders of the interest rates remain in the final specifications, with changes in the alternative yield affecting R^m with a one-quarter lag (Table 2, equations 3 and 4). The diagnostic tests reveal some heteroscedasticity (of the *ARCH* form and variously linked to the squares and cross-products of the regressors) for the $\Delta_3 R^m$ equation, and the residual normality test for the residuals from the $\Delta_3 R^b$ model is close to its 5% critical level. Figures 5 and 6 display huge breaks in 1970 and in 1974/76 for both interest rates, with some additional, though less serious, instability episodes in the second half of the 1970s and in the early 1980s for the interest rate on money R^m.

Figure 2
Time-series plots of annual inflation and interest rates

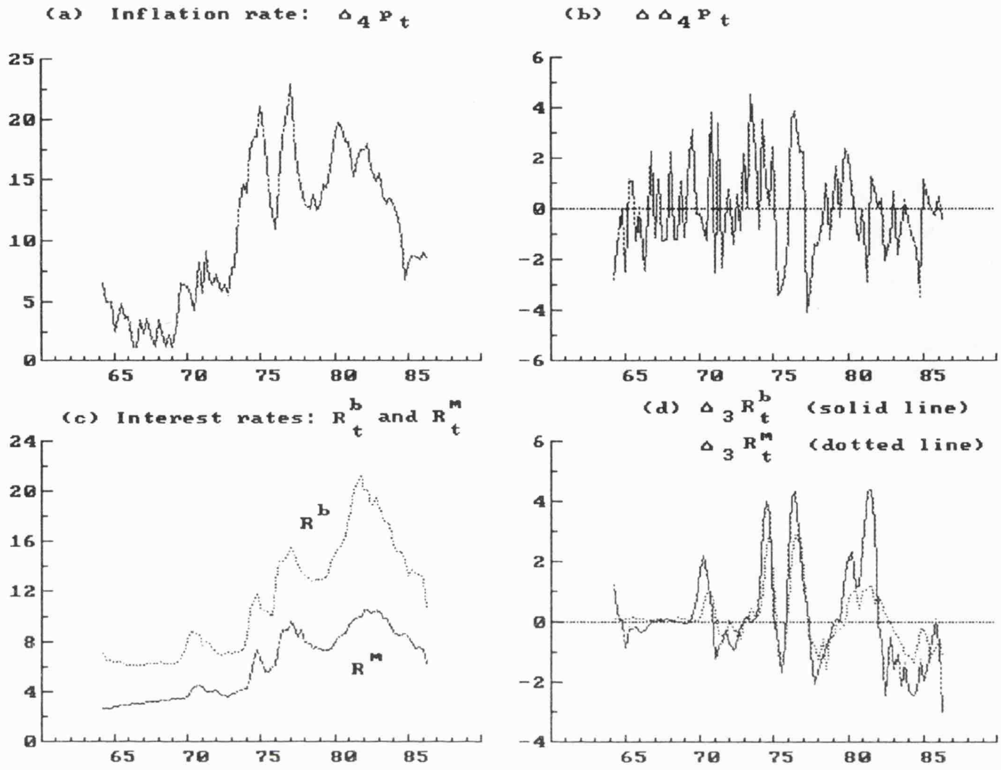


Figure 3
Stability analysis of the marginal model for $\Delta_4 y_t$

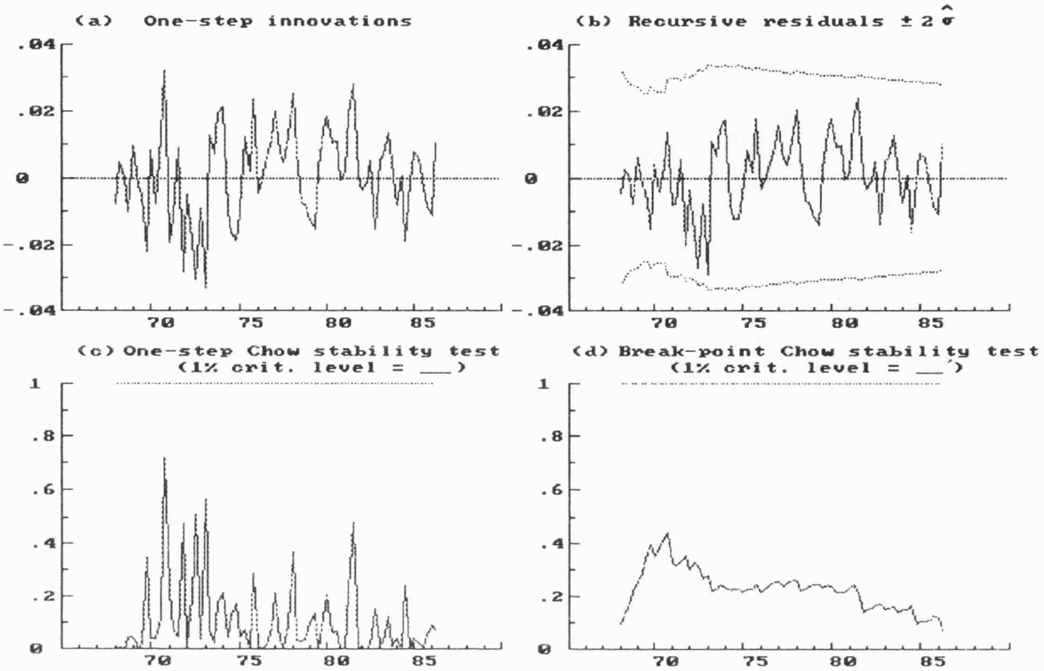


Figure 4
Stability analysis of the marginal model for $\Delta\Delta_4 p_t$

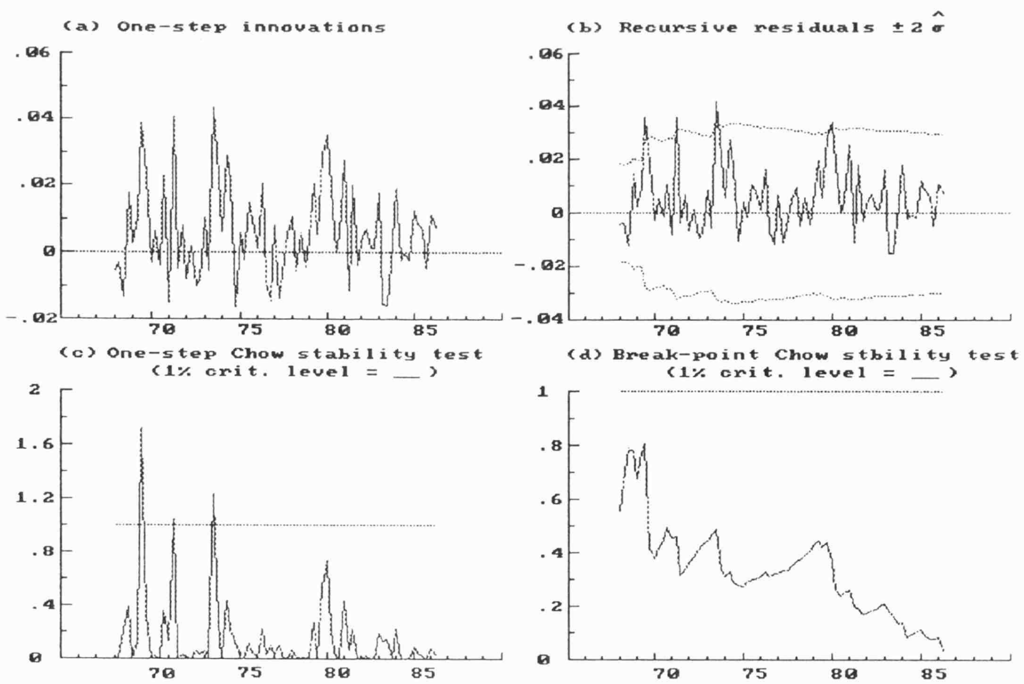


Figure 5
Stability analysis of the marginal model for $\Delta_3 R^b{}_t$

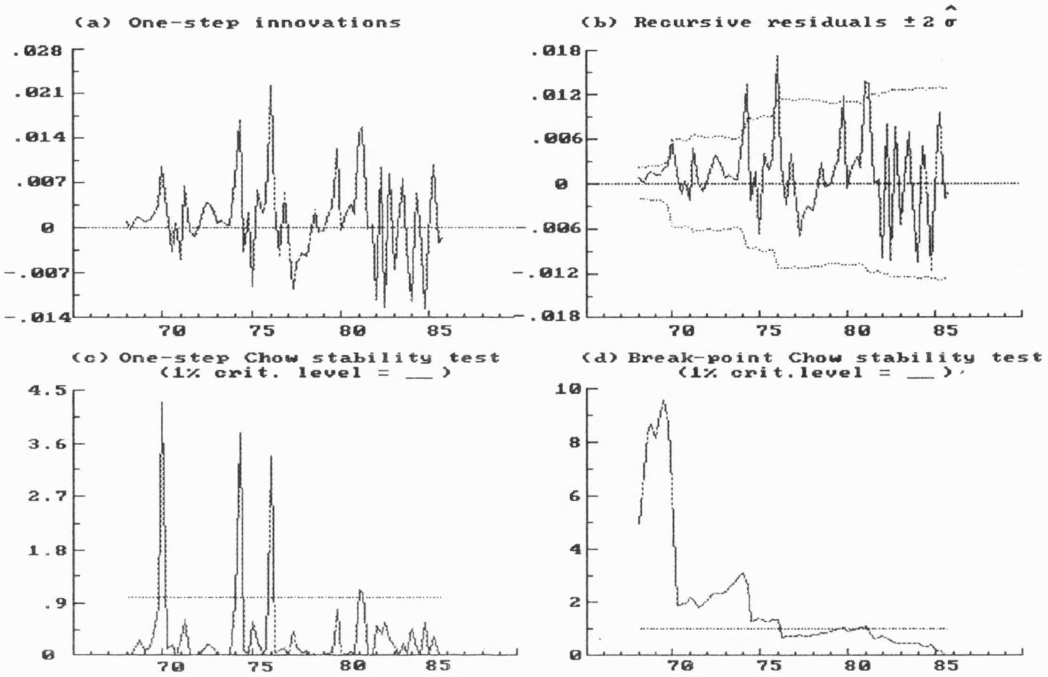
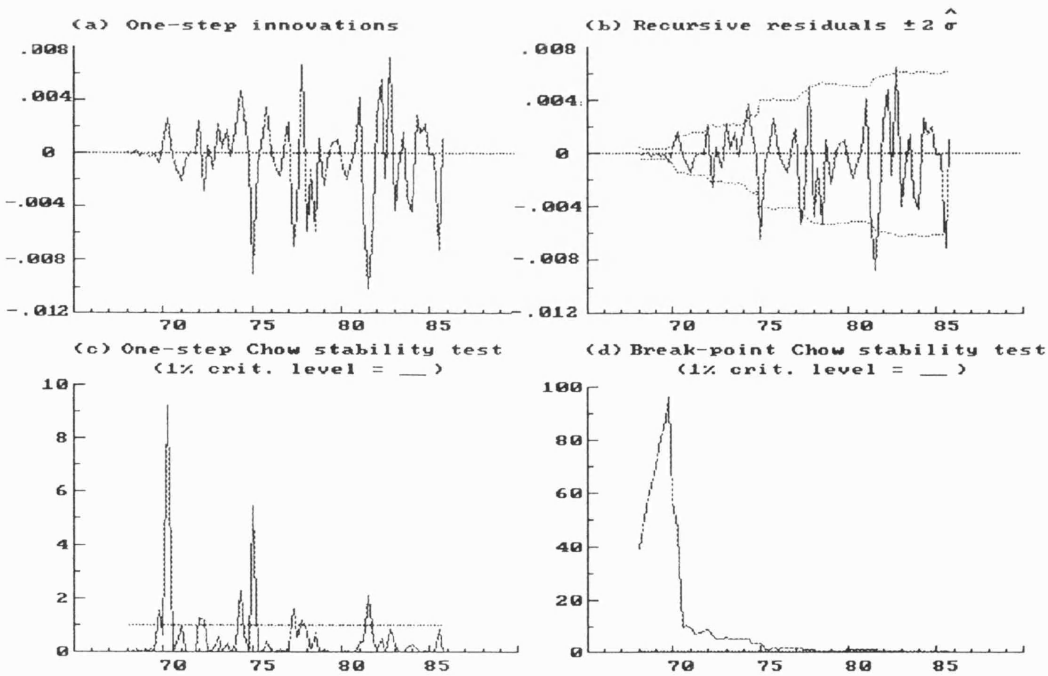


Figure 6
Stability analysis of the marginal model for $\Delta_3 R^m{}_t$



The search for stable marginal models is therefore limited to the three variables involving inflation and interest rates. Initially, in all cases the regressor coefficients of equations 2-4 in Table 2 have been allowed to vary at the dates when instability was detected. However, this specification of instability was not successful in capturing structural changes; therefore we simply used appropriately constructed dummy variables to model shifts. Even though this is an *ad hoc* procedure, in most cases some economic explanation of the breaks can be provided.

The break in the estimated process for $\Delta\Delta_4 p_t$ detected in 1969 may be caused by a surge in world commodities prices occurred at the beginning of the year; this external pressure on the inflation rate was strengthened by generalized domestic wage increases, readily transferred on prices in the second half of the year. Two other breaks are identified in 1971 and 1973, at least the latter attributable to the oil price shocks that hit western economies in that period. The introduction of point dummies for these episodes is sufficient to completely eliminate the instability from the equation. Since their estimated coefficients are of similar magnitude (capturing increases of the inflation rate ranging from 2.8 to about 4 percentage points), only one dummy variable, denoted *DINFL* and taking the value of *1* in 1969(3), 1971(2) and 1973(3) and zero everywhere else, was included in the final (stable) model reported in Table 3, equation 1. Estimated residuals, $\hat{\eta}(\Delta_4 p)$, and squared fitted values, $\hat{x}^2(\Delta_4 p)$, from this extended model will be used in the superexogeneity tests below.

The first instability episode concerning the interest rate regressors $\Delta_3 R^b$ and $\Delta_3 R^m$ at the beginning of the 1970s coincides with an important change in the conduct of monetary policy. In fact, starting from 1966, monetary authorities successfully implemented a policy of perfect stabilization of the yield on government bonds at a level of around 5.5%, chosen with reference to the conditions prevailing in international capital markets and to the domestic inflation rate, with the twofold aim of guaranteeing the absorption of large new issues of government bonds and preventing a rise in market rates during a period of slow investment activity. The remarkable increase in the demand for government bonds from the banking system and the public during this four-year period is evidence of the success of this policy also in stabilizing agents' expectations. At the end of 1969, following a generalized increase of foreign interest rates and the already mentioned surge in inflation, the central bank abandoned the interest rate stabilization policy (Ferrari (1973), Fazio (1979)). The shift towards higher interest rates was deliberately implemented by means of sharp increases of those under more direct control of the central bank, so as to reduce uncertainty over future interest rate levels (Bank of Italy (1970)). The largest increase in interest rates occurred in the first and second quarters of 1970, when the central bank finally suspended intervention

Table 3
Augmented marginal models for $\Delta\Delta_4 p$, $\Delta_3 R^b$, $\Delta_3 R^m$

1. $\Delta\Delta_4 p_t$

$$\underset{}{\Delta\Delta_4 p_t} = \underset{(0.035)}{0.071}\ \Delta_4(m-p)_{t-1} + \underset{(0.065)}{0.322}\ \Delta_4(m-p)_{t-4} - \underset{(0.057)}{0.335}\ \Delta_4(m-p)_{t-5} + \underset{(0.150)}{0.452}\ \Delta_2 R^b_{t-1}$$

$$+ \underset{(0.225)}{0.332}\ \Delta_3 R^m_{t-1} - \underset{(0.002)}{0.005} + \underset{(0.008)}{0.035}\ DINFL_t$$

	$R^2 = 0.551$	$\sigma = 1.305\%$	
$AR(5)$:	$F(5,77) = 0.54$ [0.75]	*Normality* :	$\chi^2(2) = 0.37$ [0.83]
$ARCH(4)$:	$F(4,74) = 0.81$ [0.52]	*Heterosc.*:	$F(11,70) = 1.06$ [0.40]
$X_i{*}X_j$:	$F(23,58) = 1.02$ [0.46]	*RESET* :	$F(1,81) = 1.07$ [0.30]

2. $\Delta_3 R^b_t$

$$\Delta_3 R^b_t = \underset{[0.163]}{0.319}\ \Delta R^b_{t-1} + \underset{[0.080]}{0.984}\ \Delta_2 R^b_{t-1} - \underset{[0.031]}{0.086}\ R^b_{t-1} - \underset{[0.029]}{0.045}\ \Delta_4(m-p)_{t-3}$$

$$+ \underset{[0.020]}{0.058}\ \Delta(m-p)_{t-4} + \underset{[0.037]}{0.063}\ \Delta_4 y_{t-1} + \underset{[0.022]}{0.059}\ \Delta_4 p_{t-4} + \underset{[0.025]}{0.049}\ \Delta_2\Delta_4 p_{t-3} + \underset{[0.386]}{0.264}$$

$$+ \underset{[0.183]}{0.891}\ DRb70_t + \underset{[0.306]}{1.779}\ DRb74_t$$

	$R^2 = 0.884$	$\sigma = 0.594$	
$AR(5)$:	$F(5,73) = 0.60$ [0.70]	*Normality* :	$\chi^2(2) = 11.21$ [0.01]
$ARCH(4)$:	$F(4,70) = 0.97$ [0.43]	*Heterosc.*:	$F(19,58) = 1.37$ [0.17]
$X_i{*}X_j$:	$F(49,28) = 0.95$ [0.57]	*RESET* :	$F(1,77) = 1.05$ [0.31]

3. $\Delta_3 R^m_t$

$$\Delta_3 R^m_t = \underset{[0.113]}{0.831}\ \Delta_2 R^m_{t-1} - \underset{[0.019]}{0.041}\ R^m_{t-1} + \underset{[0.086]}{0.153}\ \Delta R^b_{t-1} + \underset{[0.059]}{0.115}\ \Delta_3 R^b_{t-1}$$

$$- \underset{[0.011]}{0.028}\ \Delta_4(m-p)_{t-5} + \underset{[0.016]}{0.043}\ \Delta_4 y_{t-1} + \underset{[0.137]}{0.240} + \underset{[0.070]}{0.128}\ DRm70_{t-1} + \underset{[0.176]}{0.907}\ DRb74_{t-1}$$

	$R^2 = 0.901$	$\sigma = 0.301$	
$AR(5)$:	$F(5,75) = 0.68$ [0.64]	*Normality* :	$\chi^2(2) = 7.65$ [0.02]
$ARCH(4)$:	$F(4,72) = 2.66$ [0.04]	*Heterosc.*:	$F(14,65) = 1.42$ [0.17]
$X_i{*}X_j$:	$F(31,48) = 1.48$ [0.11]	*RESET* :	$F(1,79) = 3.65$ [0.06]

Note: The sample period is 1964(2)-1986(2). The dummy variables are defined as follows: $DINFL_t$ takes on the value of 1 in 1969(2), 1971(2), 1973(3) and 0 elsewhere; $DRb70_t$ is 1 in 1970(1) and 1970(2), -1 in 1971(1) and 0 elsewhere; $DRb74_t$ is 1 in 1974(2) and 1976(1) and 0 elsewhere; $DRm70_{t-1}$ is 1 only in 1970(2) and 1970(3) and 0 elsewhere. [.] beneath coefficient estimates denote heteroscedasticity-consistent standard errors.

to stabilize bond prices. The immediate sharp reduction in the demand for government bonds from both banks and the public suggests that this change in the monetary policy regime was clearly perceived by agents in the markets. To capture these policy-induced repeated increases in interest rates -and the following partial ease of monetary policy conditions occurred in the first months of 1971- a dummy variable, *DRb70*, taking the values of *1* in the first two quarters of 1970 and *-1* in the first quarter of 1971, is introduced in the model for $\Delta_3 R^b$; the estimated coefficient (Table 3, equation 2) indicates changes in interest rates of about 0.9 percentage points at the three relevant dates. The other break in the series for $\Delta_3 R^b$ detected in 1974/76, on the other hand, does not seem to coincide with any clear change in the prevailing policy regime, but occurs at a time of general change in the economic environment and in particular in the processes generating inflation, at least partly reflected in a new upsurge in interest rates.[12] A second dummy variable (*DRb74*) with *1* only in 1974(2) and 1976(1) is then included in the final augmented model for $\Delta_3 R^b$, which now proves stable throughout the 1970s. In addition to estimated residuals ($\hat{\eta}(R^b)$) and squared fitted values ($\hat{x}^2(R^b)$), also a four-quarter moving average of the squared residuals (denoted $\hat{\sigma}^2(R^b)$, being a measure of the possibly time-dependent residual variance) are derived from this model.

The instability for the change of the interest rate on money ($\Delta_3 R^m$) mainly mirrors the pattern detected for the yield on alternative assets with a one-quarter lag. Only the ease in the monetary policy stance in 1971(1), reflected in a sharp decline in R^b, did not affect the interest rate on money, probably because of the partial response of deposit yields in periods of declining market rates. Therefore the dummy variable included in equation 3 of Table 3 to capture instability in 1970, *DRm70*, takes the value *1* in the second and third quarters of the year, with a one-period lag relative to *DRb70* (this justifies the *t-1* time subscript on this variable). Also the lagged value of *DRb74* is able to capture mid-'70s instability in the change of money yield. As for $\Delta_3 R^b$, also in the $\Delta_3 R^m$ case a measure for the time dependent error variance of the marginal model is derived from estimation as a four-quarter moving average of squared residuals ($\hat{\sigma}^2(R^m)$).[13]

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[12] However, when the variable *DINFL*, capturing the repeated episodes of sharply accelerating inflation, is included (also with lags) in the model for $\Delta_3 R^b$ it is not statistically significant.

[13] Since the *ARCH*(4) test indicates the presence of residual autoregressive conditional heteroscedasticity, also the scaled residuals from fitting a four-order *ARCH* model are derived and used in the tests below as an alternative measure of the residual variance. The results reported in Table 4 are unchanged when this alternative measure is employed. It should also be noted that in both equations involving interest rates the outcome of normality tests is due to some outlier observations that now appear of increased importance, since the main instances of large residuals have been eliminated from the equations.

The link between the instability displayed by our feedback model (13) and the detected shifts in the marginal models for the regressors can now be formally assessed using the testing framework outlined in the preceding subsection. To this aim, according to the general formulation of the exogeneity and invariance tests in (19), the variables constructed from estimation of the extended (stable) marginal models, taken as proxies for the mean and variance of the regressors' distributions, are included in the feedback specification of money demand and the associated coefficients are tested for statistical significance. In Table 4 F-test results on individual variables and joint F-tests are reported.[14] Only changes in the mean of the alternative interest rate process (through $\hat{\eta}(R^b)$) seem to have some relevance in explaining structural changes in the feedback equation.

More informative results are obtained from the superexogeneity tests performed with the direct inclusion in the feedback model of the dummy variables used to capture regime shifts in the regressors' processes. As the lower part of Table 4 shows, the variable capturing the effect of the monetary policy regime shift occurred in 1970 onto the time-series behaviour of the interest rate R^b, *DRb70*, is highly statistically significant in the feedback equation. The determinant of interest rate non-constancy in that period seems to explain the structural shift displayed by the conditional money demand model at the same dates. A similar result is not obtained for the other episode of instability in the alternative interest rate, occurred in the mid-'70s, and for the variables modelling instability in the inflation rate and in the own-rate on money.

Overall, the results of this section suggest that the structural break of the feedback model in 1970 is exactly mirrored by a sizeable shift in the process generating interest rates, whereas no other break in the marginal models for the determinants of money demand seems to be reflected in the instability of the feedback equation. Given the particular nature of the cause of interest rate instability in 1970 -a clear and readily perceived change in the prevailing monetary policy regime- the next section explores the possibility that an explicitly forward-looking specification of money demand may be able to eliminate the instability problem at this date.

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[14] Variables pertaining to regressors involving interest rates are dated t-1 since $\Delta_3 R^b$ and $\Delta_3 R^m$ enter the feedback model (13) lagged one period. The estimation period is 1964(2)-1986(2) except when the $\hat{\sigma}^2(\cdot)$ terms are included. In these cases, allowing for four-quarter moving average construction, the estimation period begins in 1965(2).

Table 4
Invariance and superexogeneity tests on the feedback model

Variable added to feedback model	*Individual variable F-test*	*Joint variable F-test*
$\hat{\eta}(y)_t$	$F(1,76) = 0.02\ [0.88]$	$F(2,75) = 0.13\ [0.87]$
$\hat{x}^2(y)_t$	$F(1,76) = 0.27\ [0.60]$	
$\hat{\eta}(\Delta_4 p)_t$	$F(1,76) = 2.53\ [0.12]$	$F(2,75) = 1.29\ [0.28]$
$\hat{x}^2(\Delta_4 p)_t$	$F(1,76) = 0.17\ [0.68]$	
$\hat{\eta}(R^b)_{t-1}$	$F(1,75) = 5.23\ [0.02]^{*}$	$F(4,68) = 2.07\ [0.09]$
$\hat{x}^2(R^b)_{t-1}$	$F(1,75) = 0.86\ [0.35]$	
$\hat{\sigma}^2(R^b)_{t-1}$	$F(1,71) = 1.86\ [0.17]$	
$\hat{\sigma}^2 \cdot \hat{x}(R^b)_{t-1}$	$F(1,71) = 0.001\ [0.98]$	
$\hat{\eta}(R^m)_{t-1}$	$F(1,75) = 3.41\ [0.07]$	$F(4,68) = 1.17\ [0.33]$
$\hat{x}^2(R^m)_{t-1}$	$F(1,75) = 1.91\ [0.17]$	
$\hat{\sigma}^2(R^m)_{t-1}$	$F(1,71) = 1.18\ [0.28]$	
$\hat{\sigma}^2 \cdot \hat{x}(R^m)_{t-1}$	$F(1,71) = 0.39\ [0.54]$	
$DINFL_t$	$F(1,76) = 0.34\ [0.56]$	
$DRb70_{t-1}$	$F(1,76) = 9.55\ [0.003]^{**}$	$F(2,75) = 4.76\ [0.01]^{**}$
$DRb74_{t-1}$	$F(1,76) = 0.02\ [0.88]$	
$DRm70_{t-2}$	$F(1,76) = 1.79\ [0.18]$	$F(2,75) = 1.76\ [0.18]$
$DRb74_{t-2}$	$F(1,76) = 2.23\ [0.14]$	

Note: * and ** denote statistical significance at the 5% and 1% level respectively.

4. *A forward-looking alternative model.*

4.1. *The econometric specification of an expectations model for money demand.*

Although the specification of theoretical and empirical models relying heavily on agents' expectations has only recently become a deeply researched area in monetary economics, the introduction of forward-looking elements in the analysis of money demand is not a new topic in the literature. In fact, the formalization of the concept of permanent income in the late fifties led to a reconsideration of the role of current income as the appropriate scale variable in money demand functions. The ensuing debate, developed throughout the sixties, focused on the issue whether the lag structure necessary to obtain satisfactory empirical money demand equations was due to lags in the formation of expectations about permanent income (as argued by Feige (1967)), or to lags in the adjustment process of money balances to all determining variables. Subsequent research, during the seventies and early eighties, was aimed more at improving the forecasting performance of the estimated money demand equations by means of a more careful specification of the short-run dynamics, than at investigating the potential role of expectations. A notable exception is the study by Carr and Darby (1981), who formalized the notion of money as a *shock-absorber* in a simple empirically implementable form. In their model, agents form expectations on the future evolution of the (exogenously determined) money stock. Fully anticipated changes in money supply are reflected in price level expectations and therefore in nominal money demand, with no effect on real money balances. On the other hand, unanticipated monetary changes are temporarily held and, due to the sluggish movement of interest rates and the price level, do affect real balances. At the empirical level, Carr and Darby constructed a proxy for unexpected money based on a univariate time-series model and tested its significance as an additional regressor in conventional money demand equations for the U.S.. Their favourable results and the adopted estimation procedure have been subsequently challenged by MacKinnon and Milbourne (1984) on the same U.S. data, and by Cuthbertson (1986b) and Cuthbertson and Taylor (1986) on U.K. data.

More recently, a different approach, based on agents forming expectations on the determinants of money demand and not on the value of an exogenous money supply, has gained popularity. Suitable empirical formulations have been derived using the analytical framework of the multi-period adjustment cost model briefly described in section 2.1 (Cuthbertson (1988), Cuthbertson and Taylor (1987)). The loss function that agents are assumed to minimize has been described above (equation (5)) and is reported here for convenience:

$$\min_{m_{t+i}} \mathrm{E}_{t-1} \sum_{i=0}^{\infty} \phi^i \left[c(m_{t+i} - m^*_{t+i})^2 + (m_{t+i} - m_{t+i-1})^2 \right] \tag{20}$$

Expectations are formed at time *t-1* and the derived Euler equation is:

$$\phi \mathrm{E}_{t-1} m_{t+1} - (1+\phi+c)\mathrm{E}_{t-1} m_t + m_{t-1} = -c\mathrm{E}_{t-1} m^*_t \tag{21}$$

Actual money holdings are assumed to consist of a planned component, m^p, chosen to minimize (20), and an unplanned component, m^u, depending on the innovations in the determinants of money demand and capturing the buffer-stock role of money in absorbing shocks to income, interest rates and the price level:

$$m_t = m^p_t + m^u_t + \varepsilon_t \tag{22}$$

with ϵ being a zero-mean white noise stochastic process.

The solution to the above problem takes a form similar to equation (7), being:

$$m_t = \mu m_{t-1} + (1-\mu)(1-\mu\phi)\sum_{i=0}^{\infty} (\mu\phi)^i \mathrm{E}_{t-1}(p_{t+i} + \alpha_1 y_{t+i} + \alpha_2 R_{t+i} + \alpha_3 \pi_{t+i}) + m^u_t + \varepsilon_t \tag{23}$$

where use has been made of the following characterization of the target level of money holdings:

$$m^*_t - p_t = \alpha_1 y_t + \alpha_2 R_t + \alpha_3 \pi_t \tag{24}$$

Focusing on expectations on the determinants of money demand and not on money supply, as in the Carr-Darby approach, this model allows different elasticities to expected and unexpected changes in the arguments of the money demand function. However, as argued by Muscatelli (1988), only the costs of adjusting money balances and not other assets are considered, despite money being less costly to adjust than alternative assets in the buffer-stock approach. To meet this criticism, Muscatelli (1988) constructs an alternative, multiple-asset buffer-stock model in which also adjustments in non-buffer assets are penalised. The result is that the basic structure of the conventional cost of adjustment model is retained but also individuals' expectations of future saving decisions enter the determination of current money demand. If non-buffer assets are more costly to adjust, then current expected savings will appear in the equation with a positive coefficient, since agents will accumulate money holdings, which will then be gradually reallocated to alternative assets. However, the empirical results obtained from estimation on U.K. data do not seem to be sufficiently supportive for this extended buffer-stock model. Therefore, notwithstanding some unsatisfactory theoretical features, equations like (23) remain the basis of commonly formulated empirical forward-looking models of money demand and several methods have been applied to their estimation.

First, the *forward convolution method*, described, among others, by Cuthbertson (1988), hinges on the substitution for the expectations in (23) of the predictions generated by a separately estimated vector autoregression for the arguments of the target money demand. Obtaining proxies for the expected values using the *VAR* on the basis of the chain rule of forecasting presupposes that the parameters of the expectations generating processes are constant throughout the sample period. If such processes alter, fixed-parameter *VAR* yield incorrect proxies for expected values and estimation allowing time-varying parameters is needed. A second widely adopted estimation method is the *two-step procedure* proposed by Kennan (1979), based directly on the Euler equation (21). In the first step a regression of m_t onto m_{t-1} and lags of the determinants of money demand is performed in order to obtain an estimate of the stable root μ. This estimate is then used in the second step of the procedure, where actual values of m are substituted for the expectations in (21) and, after suitable transformations, the Euler equation is estimated by instrumental variable techniques, yielding estimates of the long-run elasticities α_i's.[15] The main drawback of the Kennan procedure relates to the first-step regression. Here, inconsistent estimates of the stable root μ are obtained if the lagged dependent variable m_{t-1} Granger-causes some of the arguments

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[15] In more detail, given the first-step estimate of μ and an assumption on ϕ, the adjustment cost parameter c can be calculated as $c=-(1-\mu)\phi+(1-\mu)/\mu$, using the restrictions on the sum and the product of the roots of the characteristic polynomial associated with (21). In the second step of the procedure the final specification to be estimated is obtained by transformations of the Euler equation (21). Defining the one- and two-period ahead forecasting errors as:

$$\theta_t = m_t - \mathrm{E}_{t-1} m_t = m_t^u + \varepsilon_t \quad , \quad \theta_{t+1} = m_{t+1} - \mathrm{E}_{t-1} m_{t+1} \quad , \quad \theta_t^* = m_t^* - \mathrm{E}_{t-1} m_t^*$$

and substituting the actual values of m for the expectations in (21) we get:

$$\phi m_{t+1} - (1+\phi+c) m_t + m_{t-1} = -c m_t^* + \phi \theta_{t+1} - (1+\phi+c)\theta_t + c\theta_t^*$$

Using the calculated value of c and the assumption on ϕ, the term on the left hand side of the above expression, denoted M_{t+1}, can be constructed and used as regressand. Substituting for m^*_t on the right hand side, using (24), the following estimable equation is obtained:

$$M_{t+1} = -c(p_t + \alpha_1 y_t + \alpha_2 R_t + \alpha_3 \pi_t) + V_{t+1}$$

where $V_{t+1} \equiv \phi\theta_{t+1} - (1+\phi-c)\theta_t + c\theta_t^*$. V_{t+1} contains a first-order moving average component due to the presence of θ_{t+1} and θ_t and all its terms are orthogonal to the information set at $t-1$. IV estimation is required, because of the non-zero correlation between the regressors, dated t, and the disturbance term, with instruments dated $t-1$ or earlier.

of the money demand function, being then likely to be used in the agents' expectations generating processes. The relevance of lagged money balances in the marginal models for the income and inflation variables reported in Table 2 suggests that the first step of the Kennan procedure, if applied to our data, would produce an inconsistent estimate of μ.

In the following analysis we employ an alternative -and simpler- estimation method, namely the *error-in-variables method* (*EVM*), proposed by McCallum (1976) and advocated by Cuthbertson (1990) and Cuthbertson and Taylor (1992) as an appropriate and useful procedure for forward-looking models. This method hinges on the substitution of the future expected values with the actual realizations of the same variables; the resulting equation is then estimated by instrumental variables techniques. The underlying assumption is that agents, when forming expectations, make only non-systematic forecast errors and such errors are independent of the information set on which expectations are based. We can then write, for a generic variable x: $x_{t+i}=E_{t-1}x_{t+i}+\theta_{t+i}$, with $E(\theta_{t+i}|I_{t-1})=E(\theta_{t+i}|\Phi_{t-1})=0$ for $i\geq 0$, where Φ_{t-1} is a subset of the full information set I_{t-1}. The substitution, in the equation to be estimated, of the actual realizations for the expected values determines the inclusion of θ in the error term. The resulting correlation between the regressors and the error term now requires an IV estimation technique. The properties of the forecasting error noted above imply that even if the econometrician selects the instruments from a sub-set of the information used by agents, consistent estimates of the equation parameters are derived from the IV procedure, the only requirement being that agents use at least the variables selected as instruments, a point stressed by Cuthbertson and Taylor (1992).

As a simple illustration of this property, consider the following expectations model for a variable y_t: $y_t=\alpha x_t^*+u_t$, with $E(u_t|I_{t-1})=0$. The process generating x_t is: $x_t=\delta_1 x_{t-1}+\delta_2 w_{t-1}+\epsilon_t$ and $x^e_t=E(x_t|I_{t-1})$, where $I_{t-1}=\{x_{t-1},\dots,w_{t-1},\dots\}$ represents the full information set and $E(\epsilon_t|I_{t-1})=0$. The process for x_t includes also the case of a structural change between two sub-periods, if w_{t-1} takes on zero values over the first part of the sample and becomes relevant to the determination of x only in the second sub-period. Suppose now that the set of instruments used in the IV estimation of the model for y_t -with x_t replacing x^e_t- includes only x_{t-1}. Then, when the IV method is applied to $y_t=\alpha x_t+(u_t-\alpha\epsilon_t)$, using x_{t-1} as the only instrument, consistent estimates of α are obtained, since $E[(u_t-\alpha\epsilon_t)|x_{t-1}]=0$.

As a final point, we note that the main drawback of the *EVM* is that it does not directly allow for the construction of *surprise* terms for the determinants of the demand for money, which are an important part of the forward-looking model (23), capturing the buffer- stock role of money holdings. However, proxies for such surprise terms can be constructed and will be included in the following empirical analysis.

4.2. *An alternative forward-looking model of Italian money demand.*

The empirical specification of forward-looking cost-of-adjustment models usually starts from the estimation of the unrestricted version of equation (23), with the expected values either generated from a separate *VAR* system or replaced by the actual realizations of the variables.[16] Then, the *backward-forward* restrictions derived from the declining weights structure of the coefficients on the expected variables in (23) and the presence of the stable root μ (the coefficient on the lagged dependent variable) are imposed and tested.

Our analysis of a forward-looking model for Italian money demand begins with an unconstrained specification, with the dependent variable lagged one period and the expected values (as of *t-1*) of all arguments of the demand for money function as regressors. Real money balances are chosen as the dependent variable (despite nominal balances being the obvious agents' choice variable) to make it easier the comparison of the expectations model with the purely feedback specification estimated in section 2. Moreover, from our previous analysis, the homogeneity of degree one to the price level seems to be a strong feature of nominal money balances. Estimation is performed using the error-in-variables method. The selection of appropriate instruments is carried out starting from a general reduced-form equation for *m-p* containing five lags of all variables (*m-p, y,* R^b, R^m and $\Delta_4 p$) and simplifying it to a more parsimonious model containing only statistically significant regressors. The selected instrument set includes the first, fourth and fifth lag of all variables, the second lag of both interest rates and the third lag of R^m only. The validity of the instruments will then be formally checked by means of the Sargan (1964) statistic. This test statistic is asymptotically distributed as $\chi^2(m)$ under the null hypothesis that the m overidentifying instruments are independent of the equation error. Rejection of the null hypothesis implies that some of the instruments should instead be included in the equation as additional regressors. As noted in our previous discussion, the chosen estimation method does not allow directly for the construction of surprise terms for the determinants of money demand. However, we tried to capture the unplanned component of money holdings using innovations from regressions of *y*, R^b, R^m and $\Delta_4 p$ on the instrument set. The resulting series are denoted by $Res(y)$, $Res(R^b)$, $Res(R^m)$ and $Res(\Delta_4 p)$.

In preliminary estimations, with a forecasting horizon of the length usually adopted in multi-period cost-of-adjustment models (four quarters), serious multicollinearity problems, strongly affecting the precision of the coefficients estimates, have been detected. Therefore the forecasting horizon has been limited to two periods (t and $t+1$), with the following

[16] Only if the chosen estimation method is the Kennan two-step procedure, does estimation follow a different route, outlined in section 4.1.

results (expected variables are denoted by a superscript e and replaced in estimation by actual values):

$$
\begin{array}{llllll}
(m-p)_t = & 0.858\,(m-p)_{t-1} & -\ 0.604\ y_t^e & +\ 0.886\ y_{t+1}^e & +\ 2.784\ R_t^{be} & -\ 4.332\ R_{t+1}^{be} \\
 & [0.101] & [0.560] & [0.640] & [1.436] & [1.563] \\
 & -\ 0.494\ R_t^{me} & +\ 6.143\ R_{t+1}^{me} & +\ 0.636\ \Delta_4 p_t^e & -\ 0.387\ \Delta_4 p_{t+1}^e & \\
 & [2.628] & [3.037] & [0.363] & [0.465] & \\
 & +\ 0.203\ Res(y)_t & +\ 0.982\ Res(R^b)_t & +\ 2.535\ Res(R^m)_t & -\ 1.308\ Res(\Delta_4 p)_t & \\
 & [0.260] & [1.012] & [1.527] & [0.442] & \\
 & -\ 0.134\ S_{1t} & -\ 0.145\ S_{2t} & -\ 0.168\ S_{3t} & -\ 0.376 & \\
 & [0.085] & [0.060] & [0.076] & [0.255] &
\end{array}
\qquad (25)
$$

$$\sigma = 3.30\% \qquad \textit{Specification}\ \chi^2(9) = 9.43\ [0.40]$$

Diagnostic tests [p-value]:

$AR(4)$:	$\chi^2(4) = 4.96$ [0.29]	*Normality* :	$\chi^2(2) = 2.65$ [0.27]
$ARCH(4)$:	$F(4,62) = 1.59$ [0.19]	*Heterosc.* :	$F(29,40) = 1.02$ [0.47]

Several comments on the above equation are in order. First of all, the pattern of the estimated coefficients confirms the unlikely compatibility of the data with the conventional cost-of-adjustment model of money demand already noted for Italy by Muscatelli (1991). In fact, all variables display a sign inversion, with the "right" sign on the expected values for time $t+1$, in contrast with the "declining weights" structure obtained from the theoretical model. Moreover, the coefficient standard errors are high and various regressors do not appear statistically significant. Two other notable features of the equation are related to the estimation method adopted. The high standard error of the regression, if compared with that of the feedback specification in section 2, may be partly explained by the the fact that the error-in-variables method adds to the structural residuals also the expectational errors at time t and $t+1$, contributing to the error variance. Secondly, the inclusion of expected values for time t and $t+1$ generates a first-order moving average component in the disturbance term, then justifying the possible detection of autocorrelated residuals. Although the diagnostic tests on (25) do not signal problems of this kind, the coefficient standard errors have been suitably corrected to allow for potential serial correlation (Newey and West (1987)). Finally, according to the Sargan specification test, the null hypothesis of non-correlation between the instruments and the regression residual cannot be rejected, confirming the validity of the chosen instrument set.

Overall, equation (25) marks a clear departure from the traditional multi-period cost-of-adjustment framework. When this is abandoned, no clear alternative is available for the specification of a forward-looking model of money demand. However, as also suggested by

Muscatelli (1989), a specification search on an unconstrained forward-looking equation like (25) could be conducted, in order to let the data determine the precise form in which expected values enter the equation. In such a way, one can avoid the very strong restrictions on the structure of the coefficients on the expected variables derived from the quadratic costs of adjustment model, which could be rejected by the data even when expectations play an important role in the determination of money demand. Therefore, we performed a specification search (necessarily limited, given the very short forecasting horizon) on equation (25), in order to find a more parsimonious and interpretable form for the forward-looking model.

The result from our search are reported in Table 5, where we imposed the restrictions that expected *real* interest rates, denoted by r^b and r^b, enter the model at both dates t and $t+1$, with no separate effect of the inflation rate, and only two innovations, to the own-yield on money and to inflation, are retained in the equation. Various specifications of the instrument set are tried. In column *(I)* the same instruments employed in estimating (25) are used. However, since the analysis of section 3 has shown that expectations generating processes based on these variables only suffer from instability at several dates in the sample, causing potential inference problems, other enlarged instrument sets are employed. In column *(III)* the variables introduced in the specification of the marginal models for interest rates and inflation (Table 3) to capture the main episodes of instability are added to the instruments in *(I)*. In the second column *(II)* only the two variables ($DRm70_{t-1}$ and $DRb74_{t-1}$) used to obtain stability in the time-series behaviour of the interest rate on money -equation 3, Table 3- and $DRb70_t$, capturing the break in the alternative rate due to the monetary policy shift, are added to the basic instrument set.[17] The specification test confirm the validity of the instruments used in all three cases, and the other diagnostic tests detect, at least in *(II)* and *(III)*, only some residual serial correlation (as argued above, a not surprising result, given the adopted estimation method). The standard error of the regression is in line with that of equation (13) -somewhat reduced in *(II)* and *(III)*- and the coefficient estimates, together with the statistical insignificance of the expected inflation terms when reintroduced in the equations, seem to support the imposed restrictions. The coefficient pattern is the same across the three specifications of the instrument set, with the coefficients on the interest rate variables showing a decrease in magnitude moving from the

[17] The first two variables, dated $t-1$, certainly belong to the agents' information set on which expectations are formulated. The same assumption can be justified for $DRb70_t$ on the basis of the clearly announced and readily perceived nature of the policy regime change. In any case, the estimation results are qualitatively unchanged when only the first two variables are added to the basic instrument set.

Table 5
Forward-looking specifications.

Dependent variable: $(m\text{-}p)_t$ Sample period: 1964(2)-1985(4)
Instrumental variable estimation (standard errors in parentheses)

Variable	*Instrumental variable set* I	II	III
$(m\text{-}p)_{t-1}$	0.868	0.844	0.843
	[0.093]	[0.081]	[0.087]
y_t^e	-0.882	-0.988	-1.054
	[0.334]	[0.306]	[0.310]
y_{t+1}^e	1.129	1.258	1.326
	[0.409]	[0.375]	[0.394]
$r^{b\,e}_t$	2.430	1.297	1.205
	[1.224]	[1.090]	[0.710]
$r^b{}_{t+1}^e$	-3.863	-2.560	-2.487
	[1.546]	[1.280]	[0.871]
$r^{m\,e}_t$	-2.948	-1.722	-1.600
	[1.383]	[1.161]	[0.853]
$r^m{}_{t+1}^e$	4.117	2.724	2.595
	[1.659]	[1.305]	[0.983]
$Res(R^m)_t$	2.947	2.322	2.299
	[1.134]	[1.350]	[1.360]
$Res(\Delta_4 p)_t$	-1.114	-1.041	-1.016
	[0.256]	[0.249]	[0.258]
σ	3.30%	3.06%	3.11%
Specif. χ^2	9.57 [0.57]	13.60 [0.48]	14.71 [0.55]
AR(4) $\chi^2(4)$	6.47 [0.17]	10.01 [0.04]	11.30 [0.03]
Normality $\chi^2(2)$	4.29 [0.12]	4.61 [0.10]	4.65 [0.10]
ARCH(4) *F*(4,66)	0.59 [0.67]	0.41 [0.80]	0.45 [0.77]
Heterosc. *F*(21,52)	1.29 [0.22]	1.26 [0.24]	1.24 [0.26]

Note: Estimated constant and seasonal terms not reported. The coefficient standard errors are computed following Newey and West (1987), allowing for potential residual serial correlation. The instrumental variables used in the equations are: (*I*): $(m\text{-}p)_{t-4}$, $(m\text{-}p)_{t-5}$, y_{t-1}, y_{t-4}, y_{t-5}, $R^b{}_{t-1}$, $R^b{}_{t-2}$, $R^b{}_{t-4}$, $R^b{}_{t-5}$, $R^m{}_{t-1}$, $R^m{}_{t-2}$, $R^m{}_{t-3}$, $R^m{}_{t-4}$, $R^m{}_{t-5}$, $\Delta_4 p_{t-1}$, $\Delta_4 p_{t-4}$, $\Delta_4 p_{t-5}$; (*II*): (*I*) + $DRb70_t$, $DRm70_{t-1}$, $DRb74_{t-1}$; (*III*): (*II*) + $DINFL_t$, $DRb74_t$. *Specification* χ^2 denotes the Sargan (1964) statistic for the validity of the chosen instrumental variables; the degrees of freedom, being the number of the overidentifying instruments used, are 11, 14 and 16 for specification (*I*), (*II*) and (*III*) respectively. For the tests, p-values are in [.].

basic to the enlarged instrument sets, but an increasing precision of the estimates. Overall, current money balances display a backward-looking component, captured by the lagged dependent variable, and seem to react more strongly to the values of income and (real) interest rates expected for the next period. The very similar magnitude of the coefficients on the two expected rates for $t+1$ may suggest the further restriction that the expected interest rate spread is a relevant variable in affecting money demand behaviour. Also contemporaneous unexpected realizations of inflation and of the interest rate on money have a sizeable effect on money holdings.[18]

Prior to assessing the stability properties of the equations in Table 5 and in the light of their main features (especially the short forecasting horizon and the opposite sign on the coefficients on the expected variables for time t and $t+1$) it may be interesting to briefly compare our final equation to the specification for money demand derived from the rational expectations model analyzed by Dutkowsky and Foote (1988, 1992), not adopting the cost-of-adjustment framework. In this model, an optimizing representative consumer derives utility at any date t from real consumption C_t and exchange services yielded by real money balances. Therefore, real money holdings enter the utility function directly through their role in providing liquidity services and reducing transaction costs. Each period the consumer allocates current total real income (real labour income Y_t plus interest payments on real bond holdings B_{t-1} and real money holdings $(M/P)_{t-1}$ with real interest rates r^b_t and r^m_t) to present period consumption and holdings of money and bonds. Notice that no costs of adjusting money holdings are assumed in this framework.

The consumer's optimizing problem can therefore be represented as follows:

$$\max_{C_t,\left(\frac{M}{P}\right)_t} \quad E_0 \sum_{t=0}^{\infty} \phi^t u \left[C_t, \left[\frac{M}{P}\right]_t \right] \tag{26}$$

subject to:

$$C_t + B_t + \left[\frac{M}{P}\right]_t = Y_t + (1+r_t^b)B_{t-1} + (1+r_t^m)\left[\frac{M}{P}\right]_{t-1} \tag{27}$$

The solution, using the techniques developed by Kydland and Prescott (1982) and Sargent (1989), yields a semi-reduced form for money demand of the following kind:

[18] Very similar results are obtained when the equations are respecified with $\Delta(m-p)$ as (stationary) dependent variable (imposing the restriction of a unit coefficient on $(m-p)_{t-1}$). Only the coefficients on the income terms are of closer magnitude, indicating that the growth rate of money balances is affected by the expected growth rate of income.

$$\left(\frac{M}{P}\right)_t = F\Big\{\underset{(+)}{\left(\frac{M}{P}\right)_{t-1}}, \underset{(+)}{\mathrm{E}_{t-1}r_t^b}, \underset{(-)}{\mathrm{E}_{t-1}r_t^m}, \underset{(-)}{\mathrm{E}_t r_{t+1}^b}, \underset{(+)}{\mathrm{E}_t r_{t+1}^m}, \underset{(+)}{(Y_t - \mathrm{E}_{t-1}Y_t)}, \underset{(+)}{(r_t^b - \mathrm{E}_{t-1}r_t^b)}, \underset{(+)}{(r_t^m - \mathrm{E}_{t-1}r_t^m)}\Big\} \quad (28)$$

As in the forward-looking cost-of-adjustment model, the resulting money balances equation exhibits both backward and forward looking elements. The former is represented by the lagged value of real money holdings, reflecting the effects of past realizations of income and interest rates on current money holdings and not the presence of adjustment costs. The forward-looking part of the equation distinguishes between the effects of anticipations of future variables and "surprise" terms. Speculative considerations may explain the negative effect on money holdings of current anticipations of the future yield on government bonds $E_t r^b_{t+1}$ as well as the positive effect of $E_t r^m_{t+1}$. An opposite sign pattern is found on the past anticipations of the two rates of return $E_{t-1} r^b_t$ and $E_{t-1} r^m_t$. Unexpected components of interest rates enter the equation with a positive sign, since they represent changes in interest income from bond and money holdings comparable to unanticipated variations in labour income, which also have a positive effect in the equation. Finally, note that no term reflecting the anticipated part of labour income is present in (28).

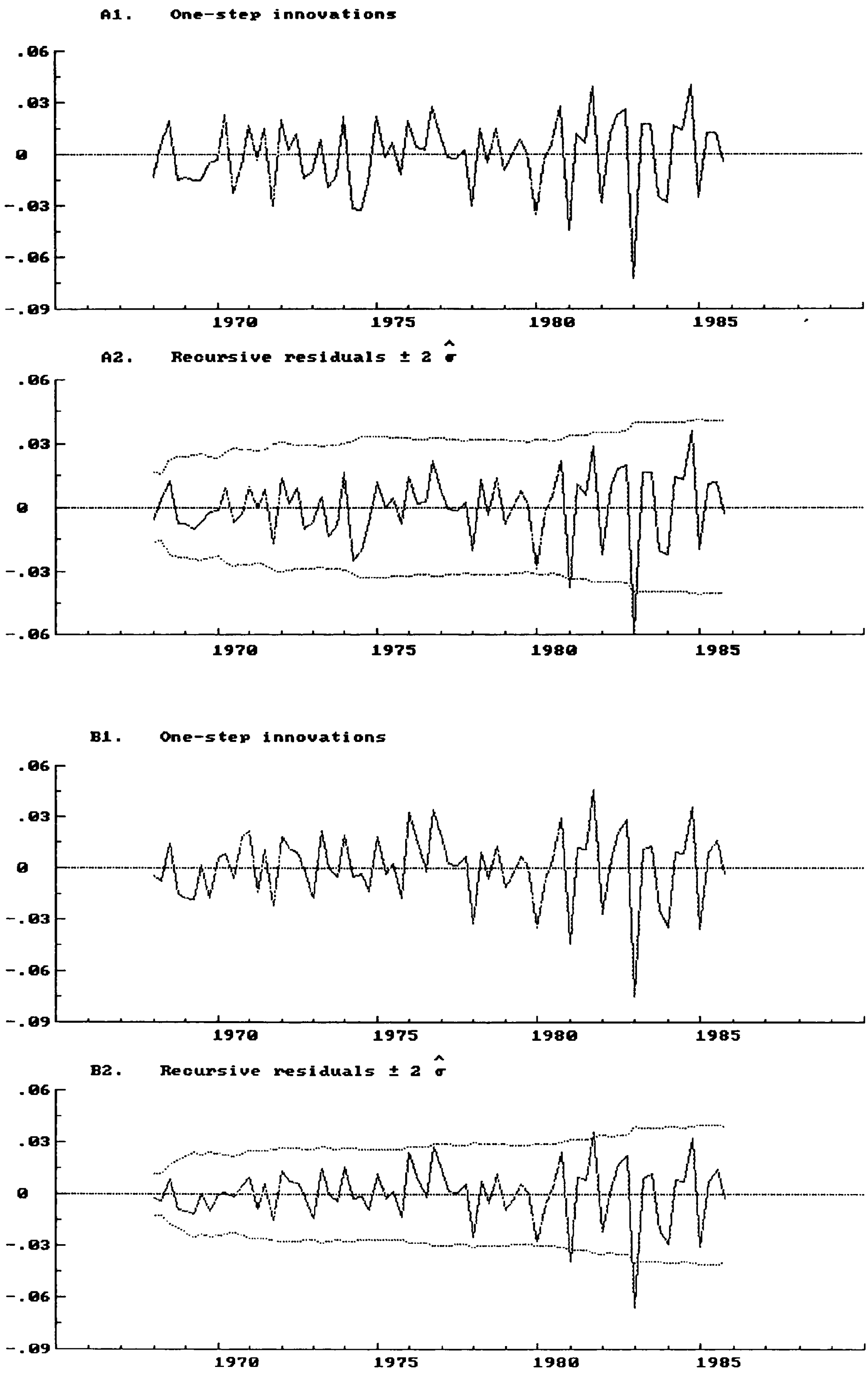
Some features of our estimated forward-looking equation closely resemble the prediction of the model by Dutkowsky and Foote. In particular, the signs of the estimated coefficients on the interest rate variables (both the expected values and the innovation terms) are in accord with those implied by the theory. However, both the separate significance of the innovation in the inflation rate and the presence of expected values of the income variable, instead of its innovation, seem difficult to reconcile with this theoretical model.[19] Furthermore, if the main concern of the analysis was to design a test of this rational expectations model for money demand, then also the implications of the theory for the behaviour of real consumption should be derived, and joint estimation of both the money demand and the consumption equations performed in order to test the implied cross-equation restrictions.[20]

The results of the stability analysis on our forward-looking model are shown in Figure 7. Since the instrumental variable estimates display fairly high residual standard errors, making it difficult the detection of instability episodes, recursive estimation has been

[19] Moreover, the income variable used in our empirical analysis (GDP) does not correspond to the labour income definition relevant to the theoretical model under discussion.

[20] Dutkowsky and Foote (1988) perform a similar analysis on U.S. data, subsequently extending the model to consider labour supply decisions (Dutkowsky and Foote (1992)).

Figure 7
Stability analysis of the forward-looking specification



performed on the second step regression of the two-stage least squares procedure applied to the restricted equation in Table 5. The standard error of this regression is around 1.6%, more comparable with that of the feedback specification. One-step innovations and recursive residuals are shown (Figure 7, panels *A* and *B*) for the instrument sets of equations *(I)* and *(III)* of Table 5. Some sign of instability is detected only in the early 1980s, with the most relevant episode in the first quarter of 1983. The equations are remarkably stable throughout the seventies, especially in the face of the monetary policy regime shift in 1970.

This last result suggests that, notwithstanding the difficulties in the theoretical interpretation of the forward-looking model estimated above, the explicit inclusion of expectations in modelling money demand may be useful to account for at least some of the structural breaks shown by feedback equations.

5. *Conclusions.*

In this chapter we compared the structural stability performance of two models of Italian money demand over the period 1964-1986: a feedback model and a forward-looking alternative, allowing for agents' expectations.

The feedback equation displays major instability at the beginning of the 1970s and in the middle of the decade. This latter episode is common to other empirical analyses of money demand behaviour in Italy and is generally attributed to some form of misspecification of the equation during a period of important financial innovations. The instability in 1970, although quantitatively more relevant, has not been stressed in previous studies. However, this very episode, coinciding with a clear change in the prevailing monetary policy regime, may be useful in assessing the potential role of expectations in determining money demand. In fact, formal tests show that the instability of the feedback specification reflects a (policy-induced) structural change in the process generating interest rates; therefore, a particular form of misspecification, namely the neglect of agents' expectations formation, becomes a candidate for explaining instability.

Although the estimated alternative forward-looking model displays some features which are difficult to rationalize, its remarkable stability throughout the seventies and especially in the face of the policy change of 1970 suggests that expectations may be an important determinant of money demand behaviour.

Appendix

A1. *Further results on the feedback specification.*

(i) When the first difference of *m-p* is chosen as dependent variable, the final feedback specification is the following:

$$\begin{aligned}\Delta(m-p)_t = & - \underset{(0.060)}{0.253}\,\Delta_4(m-p)_{t-1} + \underset{(0.049)}{0.942}\,\Delta(m-p)_{t-4} + \underset{(0.065)}{0.169}\,\Delta_4 y - \underset{(0.060)}{0.193}\,\Delta_4 y_{t-1} \\ & - \underset{(0.070)}{0.931}\,\Delta\Delta_4 p_t - \underset{(0.124)}{0.797}\,\Delta_3 R^b_{t-1} + \underset{(0.200)}{0.763}\,\Delta_3 R^m_{t-1} + \underset{(0.016)}{0.079}\,ECM^{l}_{t-1} \\ & - \underset{(0.011)}{0.053}\,D1983q4_t - \underset{(0.004)}{0.006}\,S_{1t} - \underset{(0.005)}{0.020}\,S_{2t} - \underset{(0.004)}{0.009}\,S_{3t} + \underset{(0.035)}{0.186}\end{aligned}$$

$R^2 = 0.9504 \qquad \sigma = 0.950\%$

Diagnostic tests [p-value]:

AR(5): $F(5,71) = 0.52$ [0.76]
ARCH(4): $F(4,68) = 0.21$ [0.93]
RESET: $F(1,75) = 1.71$ [0.20]
Normality :$\chi^2(2) = 6.00$ [0.05]
Heterosc. :$F(21,54) = 0.91$ [0.58]

ECM^l denotes residuals from the following long-run solution (standard errors in parentheses):

$$m - p = \underset{(0.171)}{1.582}\,y - \underset{(2.080)}{1.387}\,R^b + \underset{(4.040)}{2.645}\,R^m + \underset{(0.522)}{2.030}\,\Delta_4 p$$

Recursive stability tests yield results almost identical to those displayed in Figure 1.

(ii) Augmented Dickey-Fuller tests on the variables included in the feedback final specification (equation (13)) yield:

	Variable					
	$\Delta_4(m-p)$	$\Delta_4 y$	$\Delta\Delta_4 p$	$\Delta_3 R^b$	$\Delta_3 R^m$	*ECM*
ADF *(lags)*	-4.10** (1)	-4.03** (1)	-6.86** (3)	-4.61** (1)	-5.27** (1)	-2.65 (0)

A constant term and seasonal dummies are included in the *ADF* equation (except in the *ECM* case, where only a constant is included; in the equation for $\Delta_4(m-p)$ also a trend is added). The number of lags is chosen in order to remove residual serial correlation. * and ** denote statistical significance at the 5% and 1% level respectively. Critical values are: -2.89 (5%) and -3.50 (1%); when a trend is included: -3.46 (5%) and -4.07 (1%).

(iii) Application of the two-step procedure of Engle and Granger (1987) gives the following results:

a) estimation of the static cointegrating equation (seasonals included):

$$m-p = 1.424y - 2.931R^b + 0.984R^m + 1.450\Delta_4 p \qquad DF = -3.41^*$$

b) feedback specification with residuals from the cointegrating regression, denoted *ECMGE*:

$$\begin{aligned}\Delta_4(m-p)_t = {} & \underset{(0.031)}{1.002}\,\Delta_4(m-p)_{t-1} + \underset{(0.076)}{0.162}\,\Delta_4 y - \underset{(0.074)}{0.245}\,\Delta_4 y_{t-1} - \underset{(0.078)}{1.068}\,\Delta\Delta_4 p_t \\ & - \underset{(0.139)}{0.537}\,\Delta_3 R^b_{t-1} + \underset{(0.236)}{0.760}\,\Delta_3 R^m_{t-1} + \underset{(0.021)}{0.004}\,ECMGE_{t-1} \\ & - \underset{(0.012)}{0.062}\,D1983q4_t - \underset{(0.003)}{0.005}\,S_{1t} - \underset{(0.003)}{0.006}\,S_{2t} - \underset{(0.003)}{0.003}\,S_{3t} + \underset{(0.003)}{0.007}\end{aligned}$$

$$R^2 = 0.9701 \qquad \sigma = 1.104\%$$

Diagnostic tests [p-value]:

$AR(5)$:	$F(5,71) = 1.22$ [0.31]	*Normality* :	$\chi^2(2) = 4.22$ [0.12]
$ARCH(4)$:	$F(4,68) = 0.34$ [0.85]	*Heterosc.* :	$F(19,56) = 1.01$ [0.47]
RESET :	$F(1,75) = 0.15$ [0.70]		

(iv) Results from Johansen's (1988, 1991) procedure:

Test statistics	*Hypothesis tested* on the number of valid cointegrating vectors (r):				
	$r=0$	$r\leq 1$	$r\leq 2$	$r\leq 3$	$r\leq 4$
λ_{MAX}	32.7 (33.5)	25.2 (27.1)	13.7 (21.0)	6.4 (14.1)	3.5 (3.8)
λ_{TRACE}	81.5 (68.5)	48.8 (47.2)	23.6 (29.7)	9.9 (15.4)	3.5 (3.8)

The Johansen procedure is performed on a fifth-order *VAR* with a constant, seasonals, and the *D1983q4* dummy variable entered unrestrictedly. When the constant is included in the cointegrating space the results of the tests do not change significantly. The presence of a linear trend in the cointegrating space was tested and rejected. Diagnostic tests on the equations of the *VAR* show marked instability in the mid-'70s, as expected. No other serious misspecification problems emerge from the tests (only some sign of heteroscedasticity in the equation for ΔR^m). 95% critical values are reported in parentheses (with the small-sample correction suggested by Reimers (1992)). The estimated coefficients of the two valid cointegrating vectors (normalized on real money balances) are reported in the table below:

Estimated coefficients of valid cointegrating vectors ($r=2$) (normalized on m-p):				
m-p	y	R^b	R^m	$\Delta_4 p$
-1	1.443	-3.471	8.254	0.874
-1	1.859	1.663	-4.485	2.224

A2. *Data sources.*

The data used in the empirical analysis are obtained from the following sources:

- the components of M2 and the stock of alternative assets (government bonds, private bonds and Treasury Bills) are taken from: Banca d'Italia, Appendix to the *Statistical Bulletin*, 1972, 5, for the period 1962-1969; Banca d'Italia, Appendix to the *Statistical Bulletin*, 1983, 3-4, for the period 1970-1982; Banca d'Italia, *Statistical Bulletin*, various issues, table H1, for the period 1983-1986;

- the interest rates on M2 components, government and private bonds, and Treasury Bills are from: Banca d' Italia, *Statistical Bulletin*, various issues, tables A5 and A6;

- data on the GDP and GDP deflator are taken, for the 1962-1969 period, from: Da Empoli D., Siesto V. and Antonello P., *Finanza Pubblica e Contabilita' Nazionale Su Base Trimestrale: 1954-1975* (Quarterly National Accounts), Padova Cedam, 1979; for the 1970-1986 period from: ISTAT, Supplement to the *Monthly Statistical Bulletin*, 1983, 12, and *Monthly Statistical Bulletin*, various issues.

Chapter 3

Money demand in a multivariate framework: a system analysis of Italian money demand in the 1980s and early 1990s.

1. Introduction

As argued in chapter 1 (section 3.4), 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 (1988, 1991) and Johansen and Juselius (1990), 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 behaviour of Italian money demand, since other studies have highlighted the presence of multiple long-run relations among money balances, income and interest rates (Muscatelli (1991)), without formally testing structural hypotheses on the economic nature of the detected relations. 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 provides a detailed account of the adopted methodology, also discussing its relation with the existing literature. The empirical results are reported in section 3 and the main conclusions summarized in section 4.

2. *Methodology and related literature.*

The recent applied econometric literature focused on two main strategies for system estimation. On the one hand, following a tradition tracing back to the work of the Cowles Commission, some authors proposed the formulation 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 misspecification 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 (1994c)). On the other hand, Sims (1980) vigorously criticized the kind of exclusion restrictions commonly used for identifying structural relations and advocated the superiority of *VAR* models in capturing the complex dynamic interactions between economic variables without imposing "incredible" (over)identifying restrictions on the data: to this aim, impulse response functions and forecast error variance decompositions techniques became widely used.[1]

The non-stationary nature of most macroeconomic time series requires the adoption of appropriate methodologies for system estimation and inference. Johansen (1988, 1991)

[1] The original applications of *VAR* modelling required nevertheless some assumptions on the contemporaneous relations among *VAR* disturbances (a triangular ordering of the variables through a Choleski decomposition of the residual *VAR* matrix in Sims (1980)). More recently *structural VAR* techniques have been developed, imposing and testing theory-based restrictions on the simultaneous relations among *VAR* innovations (as in Bernanke (1986), Sims (1986), Blanchard (1989)) or long-run restrictions on the dynamic effects of the various innovations on the endogenous variables (as in Blanchard and Quah (1989)).

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 either of the above mentioned system specification strategies, in order to reach a complete characterization of the short-run dynamics of the variables, adjusting towards their equilibrium path. If Sims' approach is adopted, the usual techniques may be applied to a *VAR* including additional lagged (error-correction) terms measuring the deviations of the variables from their long-run equilibrium targets (this kind of cointegrated *VAR* analysis is applied to a small-scale macroeconomic system by King, Plosser, Stock and Watson (1991)). If, on the contrary, the alternative structural modelling strategy is followed, the cointegrated *VAR* may be 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 (1994c)).

In this chapter we follow the latter approach, combining Johansen's long-run analysis with the structural modelling strategy proposed by Hendry, Mizon and Chow, dividing our procedure into two main steps. First, we study the long-run behaviour 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. The properties of this estimated system must be consistent with the economic interpretation of the long-run equilibrium: in particular the attribution of the various disequilibrium terms to the individual equations in the system (and the estimated coefficients on these terms) must support the view that the variables react in an error-correcting fashion to deviations from the long-run equilibrium relations.

In the remainder of this section the adopted methodology is described is more detail and its connections with (and differences from) the above mentioned empirical literature briefly noted.

We begin by defining a n-dimensional kth-order *VAR* process for the vector x_t, including the non-stationary ($I(1)$) variables of interest:

$$x_t = \sum_{i=1}^{k} \Pi_i x_{t-i} + K d_t + \varepsilon_t \tag{1}$$

where d is a vector of deterministic components (constant, linear trend, seasonals) and the disturbance term vector ϵ_t is $IIN(0, \Omega)$. (1) may be rearranged to yield the following (vector) error-correction representation:

$$\Delta x_t = \Pi x_{t-1} + \sum_{i=1}^{k-1} A_i \Delta x_{t-i} + K d_t + \varepsilon_t \tag{2}$$

with

$$\Pi = \sum_{i=1}^{k} \Pi_i - I \quad , \quad A_i = -\sum_{j=i+1}^{k} \Pi_j \quad \textit{for } i = 1,\dots,k-1 \tag{3}$$

The matrix Π contains all relevant information about the long-run properties of the system.[2] Since the vector Δx_t and its lags are stationary ($I(0)$), the system in (2) displays the same degree of integration for all variables involved only if either $\Pi=0$, no level term pertaining to the right-hand side of (2), or the coefficients of Π yield stationary linear combinations of the variables in x, so that Πx_{t-1} is $I(0)$. In the latter case, the variables in x are linked by long-run (cointegrating) relations, the number of which is given by the rank of Π, r. As shown by Johansen (1988) and Johansen and Juselius (1990), three cases must be distinguished: *i)* $r=0$: Π is the null matrix and the system in (2) is a *VAR* in first differences, with no long-run relations among the levels of the variables in x; *ii)* $r=n$: Π has full rank and the vector process x is stationary; in this case a *VAR* specification in levels as in (1) is appropriate; finally, *iii)* $0<r<n$: there exist r distinct stationary linear combinations of the n $I(1)$ variables. The existence of cointegrating relations imposes cross-equation restrictions on the coefficients of the Π matrix, reducing its rank. In this last case, Π may be expressed as the product of two $n \times r$ matrices: $\Pi=\alpha\beta'$. The columns of β contain the coefficients of the r cointegrating vectors, forming the stationary combinations $\beta' x_{t-1}$, whereas the elements of α are the weights of each cointegrating relation in the equations of system (2). Johansen (1988) provides a test for the number of valid cointegrating vectors in the system together with estimates of the coefficients of α and β under the reduced rank

[2] Of course, the representation in (2) is not unique. The term in levels may well enter the system at any lag between the first and the kth, with no effect on the coefficients of the Π matrix; only the elements of the A_i matrices are affected.

assumption on Π.[3]

Correct implementation of the estimation procedure requires the disturbance term in (2), ϵ, be a normally distributed innovation. Moreover, the *VAR* system must have constant parameters. Therefore, as a preliminary step of the analysis, the estimated residuals from (2) must be tested for normality and absence of serial correlation, and stability tests must be conducted on the recursive estimates of the *VAR* parameters. At this stage, if some deviations from the assumed properties of the *VAR* is detected, modification of the chosen lag length (k in (1)) and the introduction of exogenous conditioning variables (including dummies) in the system may be used to eliminate non-normality and residual serial correlation. Moreover, exogenous and dummy variables may capture regime shifts and other episodes which are a potential cause of parameter instability. If necessary, the exogeneity status of the added variables may then be tested following Engle and Hendry (1993).

If the presence of $r>1$ cointegrating vectors is detected, the estimates of α and β delivered by Johansen's procedure cannot be immediately interpreted in terms of underlying behavioural parameters (long-run elasticities and short-run adjustment coefficients). In fact, the estimated columns of β form an arbitrary base for the r-dimensional cointegration subspace. Therefore, choosing one of the estimated vectors as a meaningful long-run relation for the economic problem at hand (e.g. money demand modelling) does not consider the possibility that this vector may well be a linear combination of (some of) the multiple equilibrium relations in the system.[4] Johansen and Juselius (1990, 1992) address this problem, providing likelihood ratio tests for the identification of the cointegrating vectors. Theory-based hypotheses on the *long-run* structural parameters may then be formally evaluated. If not rejected by the data, the long-run restrictions may be imposed on the stationary series in $\beta' x_{t-1}$, forming a vector of (restricted) error-correction terms ecm_{t-1}. When the term in levels Πx_{t-1} in (2) is replaced by Γecm_{t-1} the resulting reduced form of the system becomes a *restricted cointegrated VAR*.

Prior to formulating and testing structural hypotheses on the contemporaneous relations linking the endogenous variables and on the adjustment process to equilibrium, the

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[3] The treatment of the deterministic component in the *VAR* is important for correct inference since the presence of a constant or a linear trend in the cointegrating vectors alters the asymptotic distribution of the test statistics. In particular, when $d_t=(1,t)'$ and then $Kd_t=k_0+k_1 t$ (with k_0 and k_1 n by 1 vectors), if the constant and the linear trend are restricted to enter the cointegrating vectors, we have $k_0=\alpha\beta_0'$ and $k_1=\alpha\beta_1'$. The resulting cointegrating relations in (2) are $\beta^{*\prime} x^*_{t-1}$, where now $\beta^*=(\beta', \beta_0', \beta_1')'$ and $x^*_{t-1}=(x'_{t-1}, 1, t)'$.

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

dynamic specification of the cointegrated *VAR* may be simplified, eliminating those (lagged endogenous) variables which are not empirically relevant to the system. The resulting *parsimonious VAR* (Clements and Mizon (1991)) becomes then a suitable framework whereby simultaneous structural models may be validly tested. The specification of a parsimonious version of the (cointegrated) reduced form of the system may increase the power of the test of the overidentifying restrictions imposed on the *VAR* by the estimation of a simultaneous structural model. The final step of our methodology requires the formulation of structural hypotheses on the *short-run* dynamics of the system. In addition to the contemporaneous relations suggested by economic theory, also some hypotheses on the elements of the adjustment matrix $\mathbf{\Gamma}$, capturing the response of the endogenous variables to deviations from the equilibrium path, may be specified. The structural assumptions on the long-run behaviour of the system -tested in the previous step of the procedure- 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 on $\mathbf{\Gamma}$ (together with those on the matrix of contemporaneous relations and on the shape of the dynamics in each individual equation) may finally be tested against the system's reduced form (parsimonious *VAR*).

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. Only one long-run valid cointegrating relation is detected among the variables analyzed (real earnings, inflation, productivity, average hours worked and the unemployment rate). This vector is interpreted as a "target" relationship negatively linking real earnings (adjusted for productivity) to unemployment and included as an error-correction term in the simultaneous model, where it determines adjustment of only real earnings. Multiple cointegrating vectors are found by Hendry and Mizon (1993) and Hendry and Doornik (1994c) in estimation of a small monetary model for the U.K. (1963-1989). Here two valid long-run relations are found: one is interpreted as a demand for money function (relating real money balances to expenditure, inflation and the interest rate), whereas the other is read as an excess aggregate demand equation (linking the deviation of output from trend to inflation and the interest rate). In the structural simultaneous model[5] 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.

King, Plosser, Stock and Watson (1991) analyze the long-run properties of a three-variable macroeconomic model (estimated on US data over the period 1949-1988). The consumption-income and investment-income ratios, being stationary series, are included in the cointegrated *VAR* form of the system as valid long-run cointegrating relations. When the system is extended to include also real money balances, inflation and the interest rate, a third cointegrating vector (interpreted as a money demand function) is detected and introduced in the *VAR*. Both cointegrated systems are then subjected to the impulse response and variance decomposition analyses to assess the relative importance of permanent and transitory disturbances. No simultaneous structural model is formulated. On the contrary, studying a similar three-variable system, Chow (1993) constructs a simple structural multiplier-accelerator model, with the two ratios mentioned above capturing the long-run equilibrium of the system. 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. Here, although different sets of variables (e.g. the interest rates and various indicators of real activity) may not share a common long-run trend, the error-correction term from one group of variables (e.g. the commercial paper-Treasury bill interest rate spread) may have important explanatory power for another set of variables. The two-step procedure applied by Chow, with cointegration analysis providing the error-correction variables to be subsequently included in the simultaneous dynamic

[5] In the final estimated models simultaneity is limited to few contemporaneous relations. In the Hendry-Doornik version only the contemporaneous effect of accelerating inflation on real money holdings is included.

model, fits well into the strategy for system estimation followed by Hendry and Mizon (1993). Finally, Johansen and Juselius (1994) have recently applied a similar procedure to macroeconomic data for Australia, identifying three cointegrating vectors (an aggregate demand relation, linking the deviations of real GDP from a linear trend to real money balances, an interest rate differential, and a proxy for the long-term real bond rate). The restricted error-correction terms so constructed are then inserted in a simultaneous dynamic model for income, money, prices and interest rates. Various sets of structural hypotheses on the short-run dynamics of the system are tested and particular importance is given, throughout the identification process, to the adjustment coefficients linking the identified long-run relations to the short-run structure.

Our empirical investigation follows the spirit of 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 by 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 (albeit with a positive coefficient) 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 chapter differs from Muscatelli's analysis in at least two respects: *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. *Empirical analysis.*

3.1. *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[6]. In Table 1 the results of a battery of Augmented Dickey-Fuller (*ADF*) tests on these variables are reported. The testing strategy follows Perron (1988) and Dolado, Jenkinson and Sosvilla-Rivero (1990), starting from a general model allowing for a deterministic trend. The results 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 ($I(0)$) behaviour 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)

[6] 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". A linear model is assumed to link the observed monthly reference series to the unobserved monthly GDP series and, after appropriate transformation of the variables, estimation of the model parameters is conducted using generalized least squares methods. Application of the estimated parameters to industrial production data yields the desired estimate of monthly GDP (Barbone, Bodo and Visco (1981)).

Table 1
Integration properties of the series

Variable	Test statistics					
	t_{α^*}	t_{β^*}	Φ_3	Φ_2	t_α	Φ_1
m	-2.95	2.74	6.06	9.68**	-2.08	10.03**
p	-3.14	3.01*	5.39	5.70*	-1.25	3.69
m-p	-2.43	2.32	3.11	7.91**	-0.89	8.72**
y	-1.84	1.76	1.93	11.65**	-0.87	15.47**
R^b	-1.44	0.47	4.01	3.41	-2.79	5.00*
R^m	-0.80	0.53	3.57	4.57	-2.61	6.70**
Δm	-5.52**	-2.00	15.27**	10.19**	-5.05**	10.03**
Δp	-3.48*	-1.43	6.30	4.31	-3.21*	5.34*
$\Delta(m$-$p)$	-	-	-	-	-7.66**	29.38**
Δy	-	-	-	-	-9.67**	46.73**
ΔR^b	-	-	-	-	-4.28**	9.24**
ΔR^m	-	-	-	-	-6.50**	21.13**

Note: The test-statistics are Augmented Dickey-Fuller statistics derived, for a generic series x, from estimation of the following models:

$$(i)\quad \Delta x_t = \mu + \alpha x_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta x_{t-i} + u_t$$

$$(ii)\quad \Delta x_t = \mu^* + \beta^*(t-\frac{T}{2}) + \alpha^* x_{t-1} + \sum_{i=1}^{k} \gamma_i^* \Delta x_{t-i} + u_t^*$$

t_{α^*}, t_{β^*} and t_α are t-statistics on the estimated parameters α^*, β^* and α respectively; Φ_3, Φ_2 and Φ_1 are F-statistics for the joint hypotheses $\beta^*=\alpha^*=0$, $\mu^*=\beta^*=\alpha^*=0$ and $\mu=\alpha=0$ respectively. Critical values are tabulated in Fuller (1976, p.373) and Dickey and Fuller (1981, p. 1062-1063); statistical significance at the 5% (1%) level is denoted by * (**). The maximum lag in the estimated equations (k), chosen to obtain serially uncorrelated residuals, is: 5 for m, 4 for p, m-p and y, 3 for Δy and ΔR^b, 2 for R^b and Δm and 1 for R^m, Δp, $\Delta(m$-$p)$ and ΔR^m.

exogeneity of the inflation rate for the parameters of interest -necessary for valid estimation and inference- 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 $\Delta(m\text{-}p)$, Δy, ΔR^b and ΔR^m, in order to detect anomalies in their time-series behaviour and assess the potential role of additional exogenous variables in each individual equation. We are particularly interested in testing the residuals from such estimated reduced forms for normality and absence of serial correlation, which are necessary for the validity of the maximum likelihood procedure applied in the cointegration analysis. To this aim, letting $x_t=\{(m\text{-}p), y, R^b, R^m\}'$, we start from a basic four-lag *VAR* specification, rearranged in order to express the dependent variables in first difference form as follows:

$$\Delta x_t = A_0 x_{t-1} + \sum_{i=1}^{3} A_i \Delta x_{t-i} + \delta \Delta p_t + c + u_t \tag{4}$$

where A_0 and 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[7]), 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 2.

In all equations, huge residual non-normality is detected. For the real money balances and expenditure equations this behaviour seems attributable to isolated episodes and two dummy variables are introduced to take care of such outlier observations. In particular, in the equation for $\Delta(m\text{-}p)$, 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; Angelini *et al.*(1994) provide further information on this episode). A dummy variable (*DU878*) taking the value of 1 only in August 1987 is added to the equation for Δy, to capture a huge 5% drop in expenditure. As shown in the second panel of the table, the inclusion of these two dummies is sufficient to eliminate residual non-normality from the $\Delta(m\text{-}p)$ and Δy equations. In order to obtain a satisfactory specification for the two interest rate equations, additional dummies and also exogenous variables are needed. In the equation for ΔR^b two dummies

[7] The inclusion of Δp in (4), 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 2
Single-equation analysis

Dependent variable:	$\Delta(m\text{-}p)$	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.$\chi^2(2)$	15.08	15.20	32.22	56.96
	(0.001)	(0.001)	(0.000)	(0.000)
Ser.Corr.$F(12)$	1.47	1.43	1.90	1.18
	(0.15)	(0.17)	(0.05)	(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.$\chi^2(2)$	4.67	1.05	8.35	4.34
	(0.10)	(0.59)	(0.01)	(0.11)
Ser.Corr.$F(12)$	1.76	1.26	1.37	1.08
	(0.07)	(0.26)	(0.20)	(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.$\chi^2(2)$	-	-	1.80	1.38
			(0.41)	(0.50)
Ser.Corr.$F(12)$	-	-	1.00	1.17
			(0.45)	(0.32)
Funct.Form F	1.79	0.001	1.48	2.81
	(0.18)	(0.99)	(0.23)	(0.10)
ARCH(6) F	0.55	0.48	1.52	0.49
	(0.77)	(0.82)	(0.18)	(0.81)
Heterosc. F	1.02	3.26	0.04	0.45
	(0.31)	(0.07)	(0.84)	(0.50)
Pred. Failure $F(12)$	2.33	1.87	3.12	1.61
	(0.01)	(0.05)	(0.001)	(0.10)

Table 2/contd.

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 $\Delta(m\text{-}p)$, a dummy variable (*DUS*) is included, taking the value of 1 in December 1989 and January 1990;

ii) in the equation for Δy a dummy variable (*DU878*) is added, taking the value of 1 only in August 1987;

iii) in the equation for ΔR^b, two dummies are included. The first (*DU877*) is a point dummy in July 1987, whereas the second (*DU8967*) takes the value of 1 in June 1989 and -1 in the following month;

iv) finally, in the equation for ΔR^m, two dummies are added: the first (*DU8310*) is a point dummy in October 1983, whereas the second (*DURM3*) takes the value of 1 in three months (July 1984, September 1985 and January 1988).

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

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. In addition, so as 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. As shown in the final panel of Table 2, 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 impulses to key short-term market rates is completed within the month. 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%. 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)), estimated over the 1980-1986 period, 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, though in that case statistically different from zero.

The six dummy variables and the four additional exogenous variables discussed above are then included in system (4). 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 behaviour, responsible for most of the episodes referred to above. The system (now including dummies and exogenous variables) is 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 (1960) stability tests. On the contrary, some of them (especially the money balances and the Treasury bill rate equations) display predictive failure over 1992, as shown in Figure 2. 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 for an $F(48,80)$ distribution 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 uncertainty and the unusual riskiness of alternative financial assets perceived in that period may well be responsible for the underprediction of money balances in October (Figure 2). However, also in earlier months (notably July) 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 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.

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

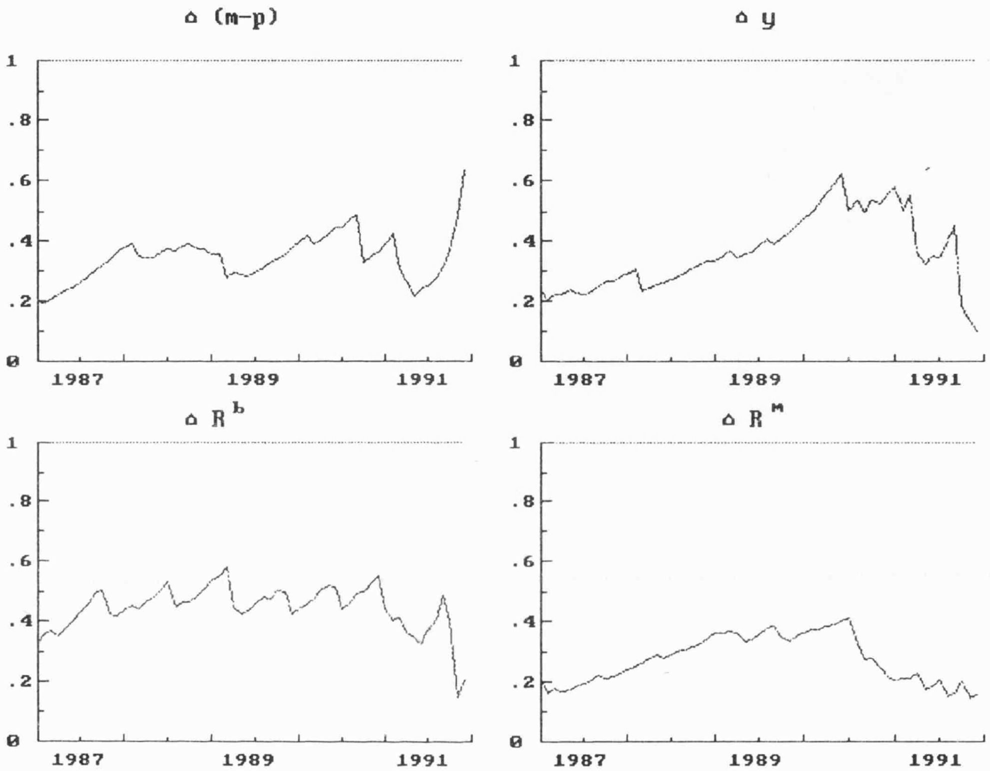
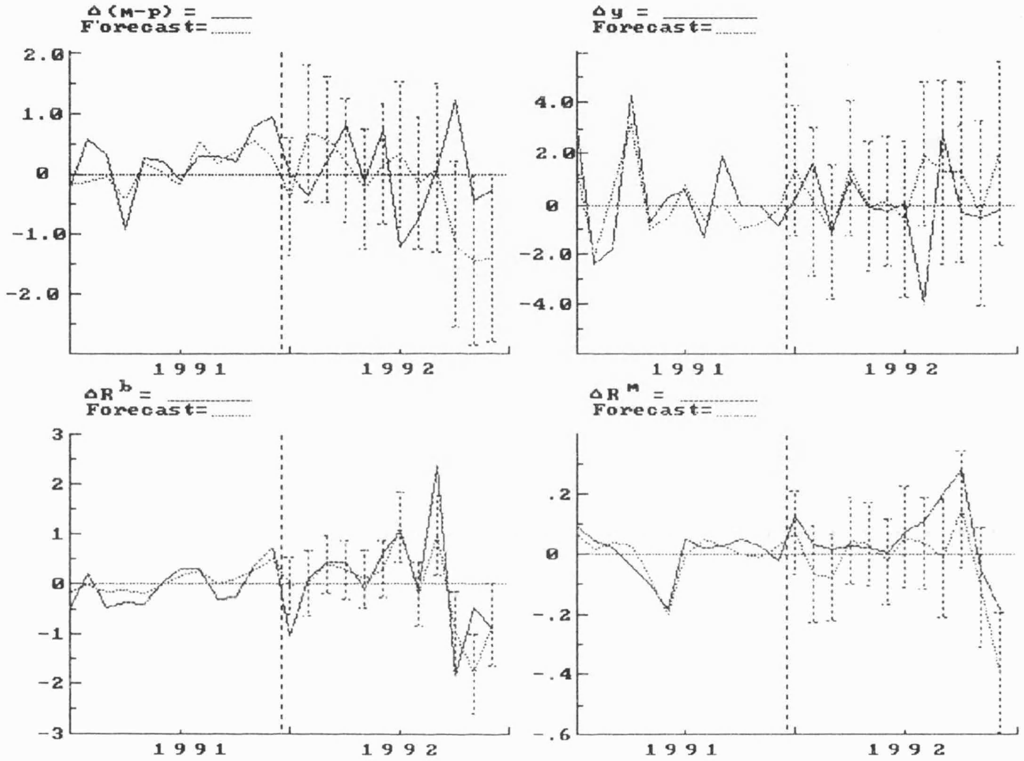


Figure 2

Forecast (with $\pm$ 2 standard error bands) from system estimation (forecast period: 1992(1)-1992(12)).



3.2. 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:

$$\Delta \boldsymbol{x}_t = \boldsymbol{\Pi} \boldsymbol{x}_{t-1} + \sum_{i=1}^{3} A_i \Delta \boldsymbol{x}_{t-i} + \boldsymbol{B}_1 \boldsymbol{d}_t + \boldsymbol{B}_2 \Delta \boldsymbol{w}_t + \boldsymbol{c} + \boldsymbol{\varepsilon}_t \qquad (5)$$

where $\boldsymbol{d}$ and $\boldsymbol{\Delta w}$ 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 $\Delta REPRN_{t-1}$) described in the preceding subsection, and $\boldsymbol{B_1}$ and $\boldsymbol{B_2}$ are conformable matrices. This four-variable system is estimated under the assumption of reduced rank of the $\boldsymbol{\Pi}$ matrix:

$$H_1(r) : \quad \boldsymbol{\Pi} = \alpha \, \beta' \qquad (6)$$

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 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 $\boldsymbol{\Delta x_t}$. Johansen's procedure allows estimation of (5) subject to the reduced rank assumption on $\boldsymbol{\Pi}$ in (6), yielding estimates of the eigenvalues of the system with corresponding eigenvectors. The *ML* procedure begins by concentrating the likelihood function of (5) with respect to the parameters in A_i, $\boldsymbol{B_1}$, $\boldsymbol{B_2}$ and $\boldsymbol{c}$ by regressing $\boldsymbol{\Delta x_t}$ and x_{t-1} onto $\boldsymbol{\Delta x_{t-i}}$ ($i=1,2,3$), $\boldsymbol{d_t}$, $\boldsymbol{\Delta w_t}$ and a constant. The residuals $\boldsymbol{R_{0t}}$ and $\boldsymbol{R_{1t}}$ are obtained and used to construct the residual product moment matrices $S_{ij}=T^{-1}\Sigma^{T}_{t=1}R_{it}R_{jt}'$ $(i,j=0,1)$. As shown by Johansen (1988), maximization of the concentrated likelihood function is obtained by solving the following eigenvalue problem:

$$| \; \lambda S_{11} - S_{10} S_{00}^{-1} S_{01} \; | = 0$$

This yields the estimated eigenvalues $\hat{\lambda}_1 > \ldots > \hat{\lambda}_n$ (n being the number of the endogenous variables in the system) and the corresponding eigenvectors $\hat{V}=(\hat{v}_1 \ldots \hat{v}_r)$, normalized such that $\hat{V}' S_{11} \hat{V} = I$. Then the *ML* estimators of α and β are given by:

$$\alpha = S_{11} \beta \qquad\qquad \beta = (\hat{v}_1 \ldots \hat{v}_r)$$

The estimated eigenvalues are then used to construct a likelihood ratio test for the number (r) of valid cointegrating vectors. Two versions of the test are available, differing in the specification of the null and alternative hypotheses. The first is based on the Maximal Eigenvalue statistic ($\lambda_{MAX}=-Tln(1-\hat{\lambda}_q)$) for testing the null $r=q-1$ against the alternative $r=q$, whereas the second is based on the Trace statistic ($\lambda_{TRACE}=-T\Sigma^n_{i=q+1}ln(1-\hat{\lambda}_i)$) for testing the null $r\leq q$ against the alternative $r\geq q+1$. Critical values for both statistics are tabulated in Johansen and Juselius (1990) and Osterwald-Lenum (1992).

The first panel of Table 3 reports the estimated eigenvalues of our four-variable system, together with the values of the λ_{MAX} and λ_{TRACE} test statistics for every possible number of cointegrating vectors. 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 ($\hat{v}_1$ and $\hat{v}_2$) 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 x_t constructed as $\hat{v}_1'x_t$ and $\hat{v}_2'x_t$ are those most correlated with the stationary part of Δx_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 $\hat{v}_1'R_{1t}$ and $\hat{v}_2'R_{1t}$, where R_{1t} has been already defined as the vector of residuals from a regression of x_{t-1} onto Δx_{t-i} ($i=1,2,3$), d_t, Δw_t and a constant. The two resulting cointegrating vectors adjusted for short-run dynamics (shown in Figures 3 and 4 normalized on m-p and R^m respectively) display the required stationary behaviour, 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 3 and 4) 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 3 cannot be given an immediate economic interpretation, since they are obtained from the estimated long-run

[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)).

Table 3

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

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

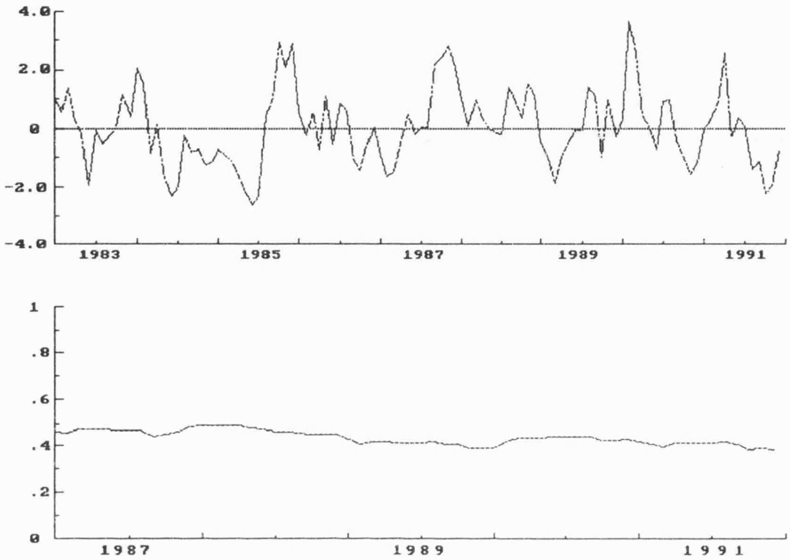
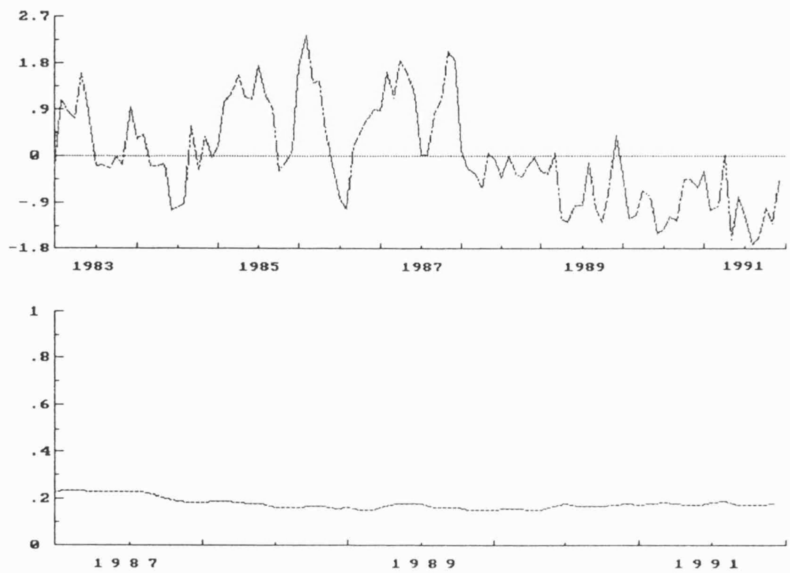


Figure 4
Residuals of the second cointegrating vector adjusted for short-run dynamics and recursive associated eigenvalue



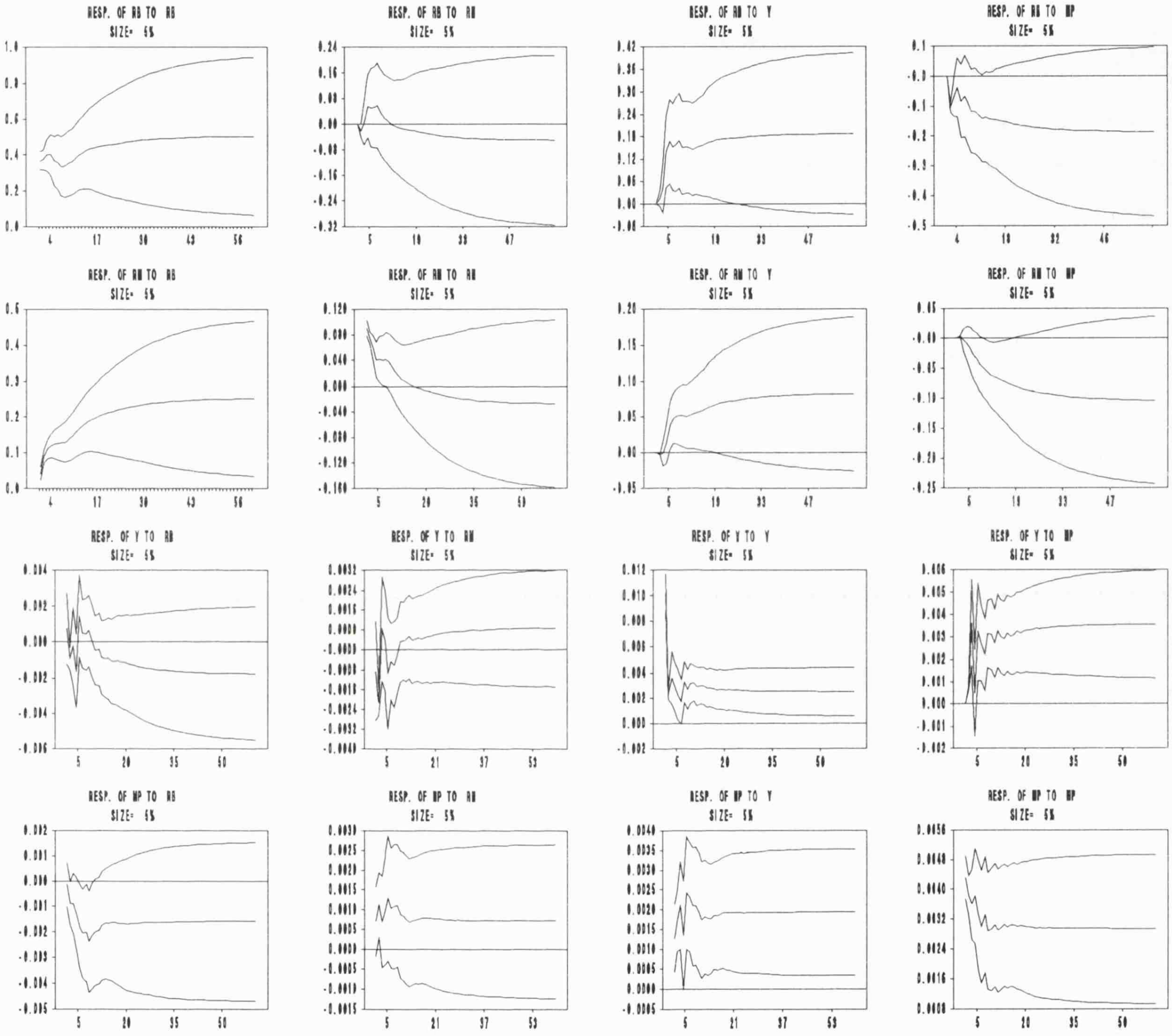
matrix Π by imposing an arbitrary normalization. Therefore, the estimated columns of β may well be linear combinations (obviously stationary) of the valid cointegrating vectors of economic interest. In order to aid economic interpretability, we present such estimates under two alternative normalizations. 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 3, 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 (5), 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.[9] The estimated impulse response functions over a sixty-month horizon are shown in Figure 5 together with 95% confidence bounds: the four columns depicts 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 5), neither money balances nor income do seem affected in a quantitatively important way.

[9] The long-run responses of the variables are qualitatively robust to changes in the ordering.

Figure 5

Impulse response functions from the cointegrated VAR.



Note: The impulse response functions are derived from simulation of the *VAR* system in equation (5) 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.

Similarly, when the disturbances have permanent effects on *m-p* and *y* (as in the last two columns of Figure 5), the two interest rates do not display in the long-run a statistically significant response.

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: $(m\text{-}p, y)$ 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 2) are included in estimation as additional stationary regressors. Moreover, in the $(m\text{-}p, y)$ -(R^m, R^b)- system, three lags of ΔR^m and ΔR^b -$\Delta(m\text{-}p)$ 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, reported in section *A1* of the *Appendix*, 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 3.

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' behaviour in setting the interest rate on deposits with reference to the bill rate, whilst the former seems not easily justifiable on the basis of available money demand theory. In fact, also 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: the pattern of spending between withdrawals as well as the amount withdrawn are exogenously fixed. However, this basic model may be generalized (as in Romer (1986) and Blanchard and Fischer (1989, ch.4)) by allowing utility-maximizing consumers to choose simultaneously the number and timing of bond 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[10]. 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[11]. 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 1980s M2 has fulfilled mainly the role of transaction medium, whereas until the late 1970s 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. The process of financial innovation occurred in the late 1970s and early 1980s 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 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

[10] The original Baumol-Tobin model was first extended to allow for wealth effects by Johnson (1970).

[11] In principle, even in the original Baumol-Tobin model money demand can be interest-inelastic (with a income-elasticity of one) 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.

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.[12]

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 totally 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.

Formally, the test is conducted by *ML* estimation of (5) subject to (6), with $r=2$, and to the following restrictions on β (henceforth, subscripts on H and the restrictions matrix M denote the specific hypothesis being tested and correspond to the panel of Table 4 where the results from estimation are reported):

$$H_{Ai}: \quad \beta = (M_{Ai}\phi , \psi) \qquad i = 1,2 \tag{7}$$

where ϕ is a *2x1* vector, containing the elements of the restricted cointegrating relation to be estimated, ψ is a *4x1* unrestricted vector and M_{Ai} denotes the following *4x2* matrices:

$$M_{A1} = \begin{bmatrix} 1 & 0 \\ 0 & 1 \\ 0 & 0 \\ 0 & 0 \end{bmatrix} \quad , \quad M_{A2} = \begin{bmatrix} 0 & 0 \\ 0 & 0 \\ 1 & 0 \\ 0 & 1 \end{bmatrix} \tag{8}$$

As shown by Johansen and Juselius (1992, section 5.3) a likelihood ratio test statistic for the above hypotheses can be constructed from the estimated eigenvalues under the restricted and

[12] 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.

unrestricted models. The test statistic is asymptotically distributed as a χ^2 variable with degrees of freedom (one in our case) given by $(n\text{-}s\text{-}r_2)r_1$, where s is the number of elements in ϕ and r_1 and r_2 are the number of restricted and unrestricted cointegrating vectors respectively. The test results (Table 4, 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 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 $a=0.880$, the estimated coefficient under A1 above, so that the cointegrating relations have the form $(z,-0.880z,*,*)$;
B2) In both cointegrating vectors the coefficients on R^m and R^b are proportional to (1,-*b*), with $b=0.495$, the value found under A2, so that the cointegrating relations have the form $(*,*,z,-0.495z)$. Again, the test is carried out by a ML estimate of the system in (5) subject to (6), with $r=2$, and

$$H_{Bi}: \quad \beta = (M_{Bi}\phi) \qquad\qquad i = 1,2 \tag{9}$$

where ϕ is the *3x2* matrix of the estimated coefficients and the restrictions matrices are:

$$M_{B1} = \begin{bmatrix} 1 & 0 & 0 \\ -a & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \end{bmatrix} \quad , \quad M_{B2} = \begin{bmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \\ 0 & 0 & -b \end{bmatrix} \tag{10}$$

with $a=0.880$ and $b=0.495$. The appropriate likelihood ratio test statistic has an asymptotic χ^2 distribution with degrees of freedom given by the number of restricted coefficients in β, two in our case (Johansen and Juselius (1992, section 5.1). The values of the 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.

Table 4
Structural restrictions on long-run relations
(* denotes imposed parameter restrictions)

A. Restrictions on *one* cointegrating vector.

A1. Two zero restrictions for coefficients on $\mathbf{R^m}$ and $\mathbf{R^b}$

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

m-p	1.943	0.481
	(-1)	(-0.208)
y	-1.710	-0.358
	(0.880)	(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	-0.152
	(-1)	(0.306)
R^b	3.894	-0.010
	(0.495)	(0.020)

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

Table 4/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	-1.517
	(-1)	(-1)
R^b	-2.099	0.774
	(0.505)	(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	-0.641
	(-1)	(-1)
y	-0.161	0.543
	(0.999)	(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)

C) Finally, we conduct a final test by assuming that *both cointegrating vectors are known*: one is proportional to *(-1,a,0,0)* with $a=0.880$ and the other is proportional to *(0,0,-1,b)* with $b=0.495$. The formal expression for this hypothesis is simply:

$$H_C: \quad \beta = M_C \tag{11}$$

with

$$M_C = \begin{bmatrix} -1 & 0 \\ a & 0 \\ 0 & -1 \\ 0 & b \end{bmatrix} \tag{12}$$

The appropriate likelihood ratio statistic (asymptotically distributed as a χ^2 with $(n\text{-}r)r_1$ degrees of freedom, r_1 being the number of the cointegrating vectors assumed known) gives a value of *4.78* with a corresponding probability value for a $\chi^2(4)$ variable 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:

$$\begin{aligned} ECMM &= (m-p) - 0.880y \\ ECMR &= R^m - 0.495R^b \end{aligned} \tag{13}$$

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.[13]

Before proceeding further, cointegration analysis is used to settle two modelling

[13] Alternative long-run hypotheses were also tested. Two of the main results, reported in the *Appendix*, section *A2*, 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 cointegtating 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.

issues raised by the chosen *VAR* specification in (5). The first is the long-run homogeneity of nominal money balances to the price level, that we imposed on the system by specifying the money variable in real terms. Given the results obtained in this section, we test price-level homogeneity in the context of a three-variable cointegrated system including separately nominal money (m), the price level (p) and expenditure (y).[14] 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 for a $\chi^2(1)$ variable 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 properties of the data. The second issue concerns the stationarity of the inflation rate detected by the *ADF* test reported in Table 1. 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.

3.3. 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 thereof) 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

[14] 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.

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 (13) 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 (5) with the two error-correction terms (lagged one period) in (13) replacing the unrestricted lagged levels of the endogenous variables. The short-run dynamics are left completely unrestricted. For this system, a *semi-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)). Table 5, panel A, reports the value of the statistics used for checking the relevant properties of the *VAR* residuals. Only those from the money balances equation display some deviations from normality, that will be eliminated below by imposing restrictions on the equation dynamics. Moreover, when residual normality, absence of serial correlation and homoscedasticity are tested at the whole system level, yielding the results reported in the last column of Table 5, no signs of mis-specification are detected.

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 ($\Delta(m\text{-}p)$, 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 section 3.1, *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.[15]

A simplification of the general dynamics of the semi-restricted *VAR* is performed by eliminating those regressors ($\Delta(m\text{-}p)_{t-2}$, 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 5, panel B. As can been seen from the high values of the corresponding F statistics, an important part of the explanatory power lies with the error-correction 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. The $F(12,209)$ test statistic for the twelve exclusion restrictions imposed by the *PVAR* onto the semi-restricted *VAR* provides formal support for the system reduction, yielding a value of 0.86, with a p-value of 0.59. Stability of the system is assessed by a recursive break-point Chow test, plotted in Figure 6 (p.119): no evidence of

[15] The values of the $F(3,92)$ statistics (and corresponding p-values) are: 1.60 (0.19), 0.39 (0.71), 1.21 (0.31) and 1.37 (0.25) for lags of $\Delta(m\text{-}p)$, Δy, ΔR^m and ΔR^b respectively; the joint $F(12,92)$ test yields a value of 0.89 (0.56). The same conclusion of absence of Granger-causality from the endogenous variables holds also for the policy rates included in the system.

Table 5

A. *Residual mis-specification tests on the semi-restricted and parsimonious VAR systems* (p-values in parentheses)

Semi-restricted VAR					
Statistic	*Equation*				*VAR*
	$\Delta(m\text{-}p)$	Δy	ΔR^m	ΔR^b	
σ	0.386	1.011	0.055	0.235	-
AR 12 $F(12,70)$	1.59 (0.11)	1.83 (0.06)	1.20 (0.30)	1.01 (0.45)	-
Normality $\chi^2(2)$	6.58 (0.04)	1.99 (0.37)	1.13 (0.57)	0.85 (0.65)	-
Heterosc. $F(38,43)$	0.31 (1.00)	0.37 (0.99)	0.96 (0.54)	0.54 (0.97)	-
ARCH(7) $F(7,68)$	0.31 (0.95)	0.20 (0.98)	0.82 (0.57)	0.75 (0.63)	-
AR(12) $F(192,126)$	-	-	-	-	1.03 (0.44)
Normality $\chi^2(8)$	-	-	-	-	9.18 (0.33)
Heterosc. $F(380,358)$	-	-	-	-	0.64 (1.00)

Parsimonious VAR					
Statistic	*Equation*				*VAR*
	$\Delta(m\text{-}p)$	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 $\chi^2(2)$	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 $\chi^2(8)$	-	-	-	-	6.74 (0.57)
Heterosc. $F(320,438)$	-	-	-	-	0.76 (0.99)

Table 5/contd.

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

$\Delta(m\text{-}p)_{t-1}$	$\Delta(m\text{-}p)_{t-3}$	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)	

F-test for the exclusion restrictions in the parsimonious *VAR*: $F(12,209) = 0.86$ [0.59]

Note: In the last column of panel A mis-specification tests are conducted at the whole system level for twelve-order serial correlation, normality and heteroscedasticity. Dummy and exogenous variables in the *VAR* are defined in notes B and C to Table 2.

structural breaks is detected.

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:

$$D_0 \Delta x_t = \Gamma \begin{bmatrix} ECMM_{t-1} \\ ECMR_{t-1} \end{bmatrix} + \sum_{i=1}^{3} D_i \Delta x_{t-i} + G_1 d_t + G_2 \Delta w_t + c + \eta_t \qquad (14)$$

where $\Delta x_t = \{\Delta(m\text{-}p)_t, \Delta y_t, \Delta R^m_t, \Delta R^b_t\}'$, d_t and Δw_t are vectors of dummy and exogenous variables respectively, and the error-correction terms *ECMM* and *ECMR* are defined in (13). The coefficients of the *4x2* matrix Γ capture the reaction of each endogenous variable to deviations from the two long-run equilibrium relations specified in section 3.2. D_0 contains the simultaneous interactions among the endogenous variables. In the process of estimating (14) various sets of identification restrictions are imposed:

a) zero-restrictions on the elements of G_1 and G_2 are imposed in order to allocate dummy and exogenous variables to the appropriate equations, according to the discussion of section 3.1 and the results of Table 2;

b) several lagged endogenous variables are excluded from the equations in which they enter with very low levels of statistical significance, therefore simplifying the dynamics described by the matrices D_i $(i=1,2,3)$;

c) most importantly, we formulate some explicit economic hypotheses on the shape of the short-run adjustment of the system to the equilibrium path by means of restrictions on the coefficients of the Γ matrix, capturing the response of the endogenous variables to deviations from the two long-run equilibrium relations. Such hypotheses are consistent with the economic interpretation of the restricted cointegrating vectors previously put forward. The lagged *ECMM* term, measuring past deviations of real money balances from a long-run relation with expenditure, is allowed to enter the equations for both $\Delta(m\text{-}p)_t$ and Δy_t. *Excess* money balances held in one period should determine an error-correcting response of the growth rate of *m-p* in the following period together with an increase of goods expenditure (a real balance effect of the sort described by the buffer-stock theory of money demand).

[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 behaviour, 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 behavioural 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 behavioural hypotheses on the long-run equilibrium and some restrictions on the dynamic adjustment of the system towards such equilibrium, also based on a behavioural interpretation.

Moreover, excess money balances may induce subsequent portfolio reallocations towards financial assets, including Treasury bills, possibly causing a decrease in the yield on such instruments: on these grounds, $ECMM_{t-1}$ is included also in the equation for ΔR^b_t. The lagged *ECMR* term, capturing deviations from the long-run equilibrium relation linking the yield on money to the bill rate, is allowed to enter the equation for ΔR^m_t (and not that for ΔR^b_t) in an error-correcting fashion. In fact, R^m is mainly determined by banks' decisions, whereas R^b is determined by market equilibrium, and therefore should plausibly display a stronger tendency to adjust towards its long-run relation with the bill rate. Furthermore, $ECMR_{t-1}$ is included in the money balances equation, so allowing for a short-run effect of temporary disequilibrium in the interest rate structure on money holdings. Given the above restrictions, the estimated matrix $\mathbf{\Gamma}$ is of the following form:

$$\mathbf{\Gamma} = \begin{bmatrix} \gamma_{11} & \gamma_{12} \\ \gamma_{21} & 0 \\ 0 & \gamma_{32} \\ \gamma_{41} & 0 \end{bmatrix} \tag{15}$$

In the initial estimation of the system (by Full Information Maximum Likelihood), the matrix of the simultaneous relations D_0 is left unrestricted and identification is achieved by imposing the restrictions under *a*) and *c*) above. In so doing, although the monthly frequency of the data tends to reduce simultaneity (all correlations between *PVAR* residuals are below 0.3, in only one case reaching 0.2), we let the data determine which contemporaneous relations are important. Then, based on the results of this initial estimation, a specification search is conducted on each equation, restricting the dynamics as mentioned above under *b*) and retaining only the statistically significant contemporaneous relations among the endogenous variables. The overidentifying restrictions so imposed in each successive step of the reduction process are evaluated by means of Likelihood Ratio tests and not rejected by the data.

This procedure leads to the final specification of the simultaneous model reported in Table 6, panel A. All equations display very simple dynamics and, as expected, few simultaneous relations are statistically significant. In particular, the equation for real balances shows a dynamics shaped only by lagged rates of change of expenditure and movements in the own return on money. The latter variable has also a contemporaneous positive effect, the most significant simultaneous relation in the whole system. The bill rate and expenditure are found to have contemporaneous interactions, with increases of R^b affecting negatively

Table 6
Simultaneous model (FIML estimation)

A. *Coefficient estimates (standard errors)*

Equation for:

Variable:	$\Delta(m\text{-}p)_t$	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	0.312	-	-0.057
	(0.035)	(0.075)		(0.020)
$ECMR_{t-1}$	0.893	-	-0.082	-
	(0.139)		(0.013)	
$\Delta(m\text{-}p)_{t-1}$	-	-0.414	-	-0.091
		(0.188)		(0.048)
$\Delta(m\text{-}p)_{t-3}$	-	-0.314	-0.016	0.151
		(0.194)	(0.010)	(0.050)
Δy_{t-1}	-0.073	-0.494	-	-
	(0.030)	(0.086)		
Δy_{t-2}	-	-0.192	-	-
		(0.074)		
Δy_{t-3}	-0.062	-	-	-
	(0.024)			
ΔR^m_{t-1}	0.620	-	0.165	-
	(0.373)		(0.052)	
ΔR^m_{t-2}	-	-	-0.076	-0.517
			(0.052)	(0.221)
ΔR^m_{t-3}	-	0.380	-	0.379
		(0.080)		(0.242)
ΔR^b_{t-3}	-	-0.586	-0.037	-
		(0.259)	(0.015)	
Δp_t	-1.088	-1.526	0.093	-
	(0.214)	(0.483)	(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)	

Table 6/contd.

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

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

Statistic	*Equation*				*Model*
	$\Delta(m\text{-}p)$	Δy	ΔR^m	ΔR^b	
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)	2.63 (0.27)	-
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 6
Break-point Chow stability test from recursive PVAR system estimation: 1987-1991
(1.0 = 5% crit. value of the test)

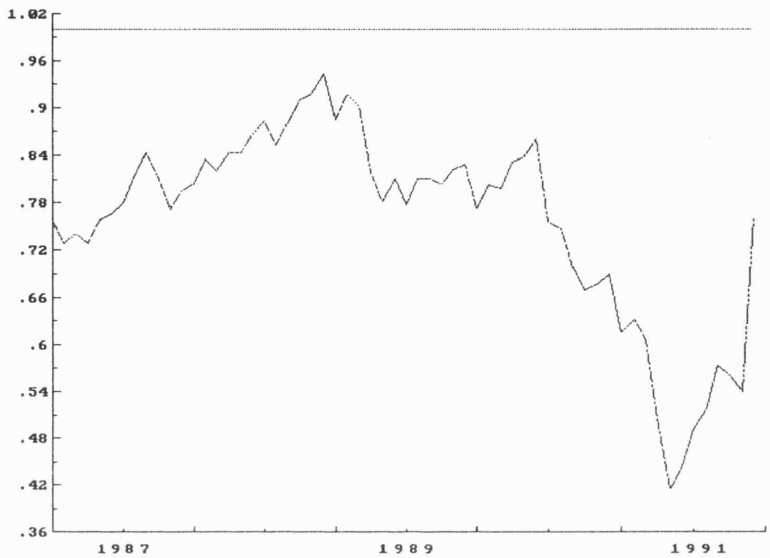
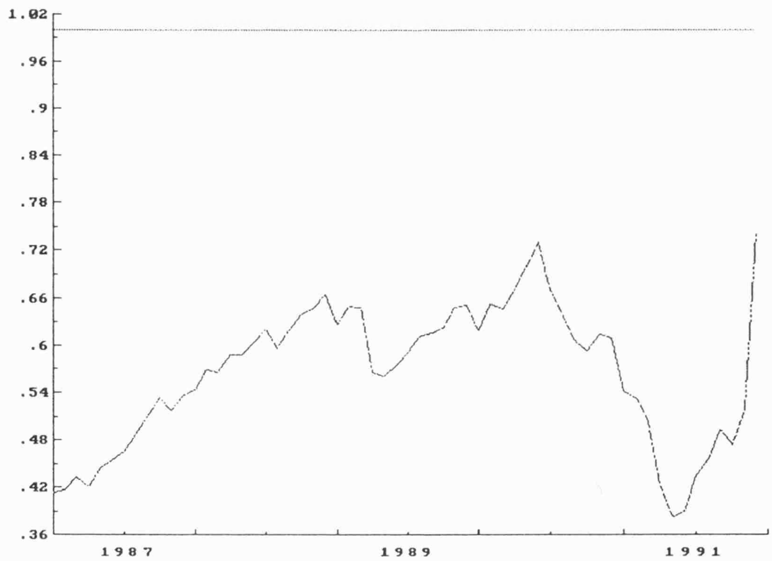


Figure 7
Break-point Chow stability test from recursive simultaneous model estimation: 1987-1991
(1.0 = 5% crit. value of the test)



expenditure and the growth of y determining a (relatively small) increase in the bill rate. The estimated coefficients on the error-correction terms are in accordance with the economic interpretation of the cointegrating vectors put forward above. Excess money (*ECMM*) enters the equation for real balances with a negative sign, positively affects expenditure on goods, and also causes a slight drop of the bill rate (this effect is small -a one per cent deviation from equilibrium balances is followed by a decrease of the bill rate by 6 basis points- but is statistically sufficiently well determined to sustain the above economic interpretation). Positive deviations of R^m from the long-run relation with the bill rate (measured by the *ECMR* term) induce error-correcting responses of the own yield of money and affect also the short-run behaviour of real balances, causing a (temporaneous) increase in money holdings. The test results reported in panel B of Table 6 do not indicate any mis-specifications of either the individual equations or the system as a whole; also the residual serial correlation previously detected in the *PVAR* equation for expenditure has been removed. The simultaneous model has a total of 50 restrictions imposed on the *PVAR* system and the likelihood ratio statistic (with a $\chi^2(50)$ distribution) for testing their validity yields a value of 46.9 (p-value 0.60), confirming that the data do not reject the model's final specification. Break-point Chow stability tests conducted on the recursive estimates of the simultaneous model (Figure 7) 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:

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

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.[18]

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.

[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 $\Delta(m\text{-}p)$, ΔR^m and ΔR^b, and with the dummy variable *DU878*, the results are unchanged.

[18] As in the multivariate analysis, the hypothesis of a long-run unit elasticity of real balances to expenditure is not rejected, yielding $\chi^2(1)=2.61$ (0.11), whereas the interest rate differential restriction is strongly rejected, with a $\chi^2(1)=7.09$ (0.008).

4. *Conclusions.*

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 behaviour 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.

Appendix

This Appendix provides further results completing the cointegration analysis of section 3.2 in the text. In section A1 the two-variable systems (*m-p, y*) and (R^m, R^b) are separately analysed. In section A2 additional structural hypotheses on the long-run relations in the complete four-variable system are tested.

A1. *Cointegration analysis of two-variable systems.*

The Johansen (1988, 1991) procedure is applied here to the bivariate systems including, as endogenous variables, real money balances and income (*m-p, y*), and the two interest rates (R^m, R^b) respectively. In the money-income system, beside Δp and the dummy variables pertaining to the money balances and income equations (*DUS* and *DU878*), also three lags of ΔR^m and ΔR^b are included as additional *I(0)* variables, to allow for more general short-run dynamic effects. Similarly, in the interest rates system, beside inflation, also the policy rate variables ($\Delta DISC_t$, $\Delta DISCN_{t-1}$, $\Delta REPR_t$ and $\Delta REPRN_{t-1}$), the dummy variables belonging to the interest rate equations (*DU8310, DURM3, DU877* and *DU8967*), and three lags of $\Delta(m-p)$ and Δy are included as additional stationary variables.

More general specifications of the cointegrating vectors are adopted. In the (*m-p, y*) system a linear trend is restricted to appear in the long-run relation and a formal test on the associated coefficient is then performed. In the (R^m, R^b) system, the presence of a constant in the cointegrating vector is tested.

	System: (*m-p, y*)		(R^m, R^b)	
Eigenvalues:	0.196	0.105	0.162	0.060
Hypothesis:	r=0	r≤1	r=0	r≤1
λ_{MAX}	23.6	12.0	19.1	6.8
95% crit. value	19.0	12.2	15.7	9.2
λ_{TRACE}	38.6	12.0	25.9	6.8
95% crit. value	25.3	12.2	20.0	9.2

Estimated valid cointegrating vector:

(normalized on *m-p*)		(normalized on R^m)	
m-p	-1	R^m	-1
y	0.734	R^b	0.501
trend	0.004	*const.*	0.683

Estimated adjustment coefficients:

(normalized on *m-p*)		(normalized on R^m)	
m-p	0.137	R^m	0.070
y	-0.493	R^b	0.063

LR test of restrictions (p-value):

coeff. on *trend* = 0	*constant* = 0
$\chi^2(1)$ = 0.004 [0.95]	$\chi^2(1)$ = 1.03 [0.31]

Figures A1 and A2 show the residuals from the estimated cointegrating vectors in the (*m-p*, *y*) and (R^m, R^b) systems respectively, with no trend or constant terms, and adjusted for short-run dynamics. The associated eigenvalues from recursive estimation of the systems are also plotted (over the 1987-1991 period) in order to assess their constancy. The vector in Figure A1 (A2) is normalized on *m-p* (R^m).

A2. *Tests of additional long-run structural hypotheses.*

Hypothesis tested in system (m-p, y, R^m, R^b):	*LR* χ^2 [*p*-value]	Comments
Fixed cointegrating vector: (-1, 1, 0, 0)	0.43 [0.51]	Hypothesis that velocity is a valid cointegrating relation not rejected.
(coeff. on *m-p*) = -(coeff. on *y*) in both cointegrating vectors	0.87 [0.65]	Velocity restriction not rejected. The estimated coefficients on R^b, normalized on R^m, are 0.48 and 0.60.
Fixed cointegrating vector: (0, 0, -1, 1)	13.79 [0.0002]	Hypothesis that the interest rate differential is a valid cointegrating relation strongly rejected.

Figure A1
Residuals from cointegrating vector in (m-p, y) system and associated recursive eigenvalue

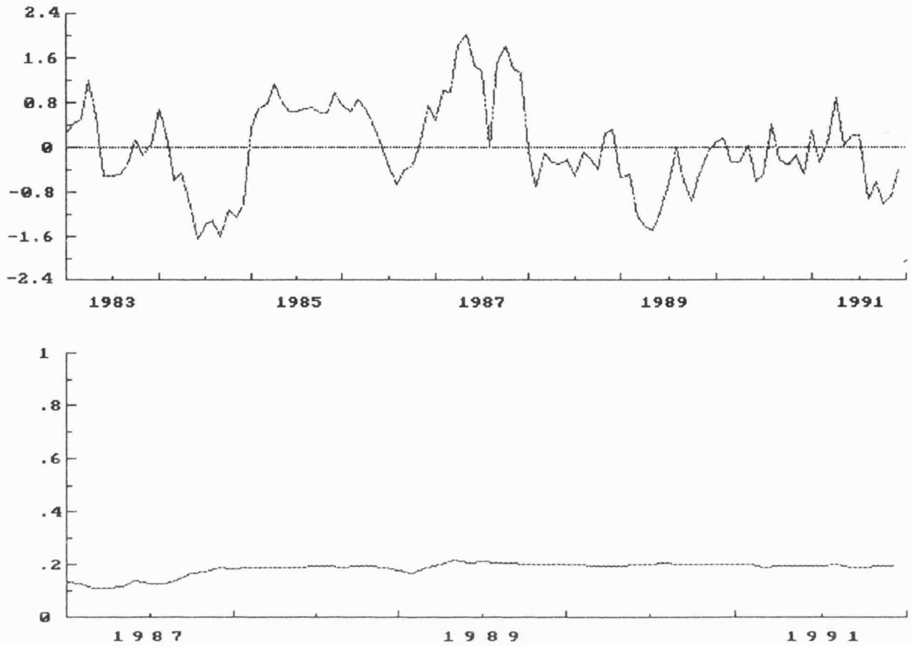
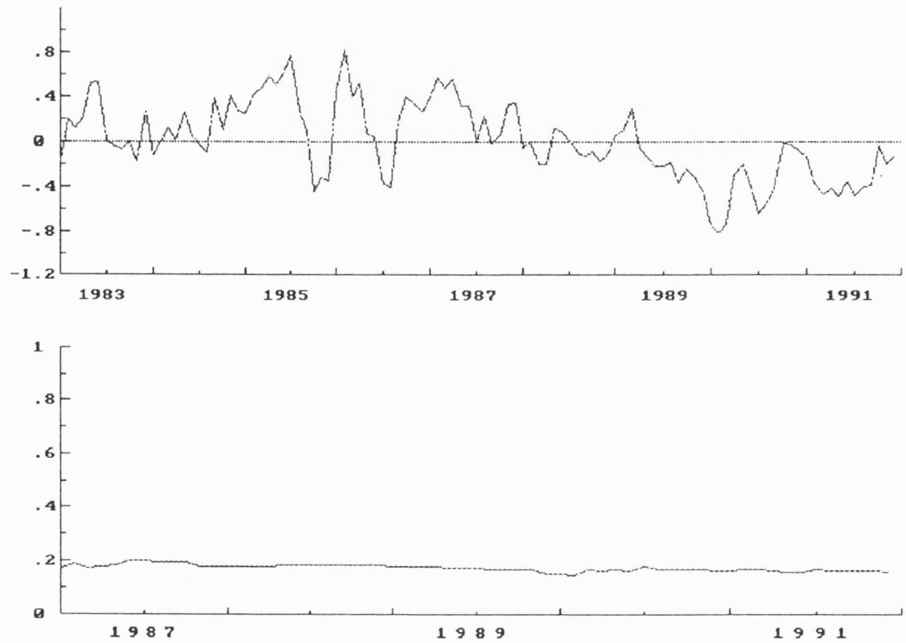


Figure A2
Residuals from cointegrating vector in (R^m, R^b) system and associated recursive eigenvalue.



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Part II

Chapter 4

Active monetary policy and the real effects of nominal shocks. A comparison of the interwar and postwar U.S. experience.

1. Introduction.

Recent developments in macroeconomics have gradually downplayed the importance of monetary factors in generating and shaping business cycle fluctuations. Real Business Cycle (*RBC*) theories, focusing on the role of preferences and production possibilities in endogenously determining real variable dynamics, completely abstract, at least in their extreme form, from monetary phenomena. A negligible role is attributed to monetary disturbances as impulses generating business cycles and the relevance of monetary policy in determining the amplitude and length of fluctuations is denied. The observed comovements of nominal and real variables are then explained mainly by a reverse causation argument, with monetary aggregates endogenously responding to output movements (King and Plosser (1984)).

At the theoretical level, various attempts have been made at explicitly considering a non-trivial role for money in an *RBC* framework (Eichenbaum and Singleton (1986), Cooley and Hansen (1989), Kydland (1989); Stadler (1994) provides a recent assessment of this strand of literature). However, none of these approaches yields a quantitatively important role of money in business fluctuations and the results of empirical investigations of this issue are viewed as broadly consistent with the predictions of *RBC* theories. In particular, in the U.S. case, causality tests often show that money has no additional predictive power for output and, most importantly, innovation accounting exercises conducted in the context of small-scale vector autoregressive systems do not detect any sizeable influence of monetary shocks on real variables.

The main aim of the present chapter is to provide an example demonstrating that such empirical findings can be generated also by a *monetary* (as opposed to a strictly *real*) equilibrium model of the cycle, when (monetary) stabilization policy is appropriately taken into account. The basic model employed is an extension of Lucas (1973) original work, modified by the introduction of an expected inflation (Mundell-Tobin) effect on aggregate demand, providing a channel for stabilization policy effectiveness.

Our results show that some evidence apparently supporting *RBC* theories may be consistent also with an extended theoretical apparatus where monetary factors are an important source of cyclical variability and monetary policy, implemented on the basis of feedback rules, is extremely effective. The implication is that the empirical findings of the kind mentioned above cannot conclusively discriminate among alternative theories of the cycle. However, the comparative analysis of time periods characterized by different policy regimes with a varying degree of stabilization effort might yield more reliable information on this issue, since under a pure *RBC* view changes in the prevailing monetary policy regime should not have any important effect on the output response to nominal shocks, whereas in a "monetary" model with a powerful stabilizing role for policy such reaction could be importantly affected. The empirical application of this chapter follows this lead for the U.S..

The scheme of the chapter is as follows. Section 2 reviews the theoretical position of *RBC* theories on the relevance of monetary factors in determining output fluctuations and the kinds of empirical evidence usually presented in support of the *RBC* view. Section 3 describes and solves the modified Lucas model, taken as the representative monetary equilibrium model of the cycle. The implications of alternative monetary regimes are then derived and discussed. Section 4 provides an empirical test of the model on U.S. data, comparing the relevance of nominal (monetary) disturbances in determining output variability in two periods (1922-1940 and 1952-1968) characterized by a different policy stance, with monetary policy systematically used for stabilization purposes only in the postwar years. Section 5 summarizes the main conclusions.

2. *The Real Business Cycle view of money: theory and evidence.*

The *basic RBC* model focuses on those properties of preferences and technology that can endogenously generate dynamics of aggregate variables and abstracts completely from monetary factors and any kind of informational imperfection. Most original contributions in the *RBC* literature, notably those of Kydland and Prescott (1982), Long and Plosser (1983), Prescott (1986), King, Plosser and Rebelo (1988a,b) and Plosser (1989), do not contain any consideration for monetary phenomena. The essential purpose of this class of models is to show how agents' optimal choice of consumption and production can display some of the main features of actual business cycles. The emphasis is on the persistence (serial correlation) and comovements (cross correlations) of the deviations from trend in aggregate variables like output, consumption, and employment.

This feature has immediately attracted radical criticisms by authors who strongly support the view that "money matters" and that any model abstracting from the characteristics of the (monetary) exchange mechanism cannot provide an adequate description of the functioning of the economic system (Summers (1986), Mankiw (1989)). However, the absence of money from the basic *RBC* model has been subjected to different interpretations. Two of them are put forward by Eichenbaum and Singleton (1986). The first (*stronger*) states that in *RBC* models "*monetary institutions and monetary policy are assumed to be inherently neutral*" (p.91), with monetary policy actions and the characteristics of financial intermediaries incapable of affecting real allocations. The second (*weaker*) interpretation is that "*the market organizations and the nature of monetary policy in the sample period being examined* [post World War II in the U.S.] *are such as that an RBC model provides an accurate characterization of the real economy*" (p.91-92). Accepting the latter interpretation, Eichenbaum and Singleton conclude that:

> "*proponents of RBC theories are not claiming that monetary policy cannot or has never had a significant impact on the fluctuations of real output, investment, or consumption. Rather, we subscribe to the second interpretation of RBC analyses as investigations of real allocations under the assumption that, to a good approximation, monetary policy shocks have played an insignificant role in determining the behaviour of real variables*" (p.92)

and that monetary policy rules have not had an important role in stabilizing the economy in the face of (non-monetary) exogenous shocks. The existence of a stronger and a weaker interpretations of the *RBC* view of monetary phenomena is confirmed also by McCallum (1989):

> "*It is not true ... that* [*RBC*] *models must be interpreted as implying the literal absence of money. Indeed, it is doubtful that RBC proponents intend to advance the proposition that no less output would be produced in the United States (with the existing capital stock) if there were no medium of exchange -that is, if all transactions had to be carried out by crude or sophisticated barter. But* [*RBC*] *models do imply that, to a good approximation, policy-induced fluctuations in monetary variables have no effect on real variables..., at least for fluctuations of the magnitude experienced since World War II*" (p.34).

The basic *RBC* models have been extended in various ways to allow for monetary factors and financial intermediation, thereby accounting for the actual comovements between monetary (or, more generally, nominal) and real variables over the business cycle. On the one hand, King and Plosser (1984), following the approach of Fama (1980) and Fischer (1983), view the provision of transactions and accounting services as the essential function performed by the financial (banking) system and model banks simply as producers of a particular intermediate good. Eichenbaum and Singleton (1986) and Cooley and Hansen (1989) also stress the role of money as a means of exchange but adopt the approach of Lucas (1980) and Lucas and Stokey (1983), introducing money in an equilibrium business cycle model by means of a *cash-in-advance* constraint.

An alternative approach is followed by Williamson (1987), who constructs an equilibrium model of the cycle in which financial intermediation plays an essential role in funding investment projects that could not be financed directly on the capital market because of asymmetric information and large costs in monitoring investors' performance. In Williamson's model real shocks are capable of generating many of the observed comovements among nominal and real variables and in particular a positive correlation between the unexpected component of movements in the price level and real activity. On the other hand, shocks to the money supply determine comovements in aggregate time series which are inconsistent with the observed cyclical pattern. The main conclusion is therefore that *real* theories of the cycle are more satisfactory explanations of the actual behaviour of the economy than traditional *monetary* theories. Williamson's model is close in spirit to a rapidly growing body of literature trying to understand the role of financial intermediation in propagating the effects of shocks to the economy, and to assess the relevance of shocks to the financial system itself (changes in regulations, innovations and technical progress in the intermediation process) as one of the driving forces of the cycle (Bernanke and Gertler (1987), Gertler (1988)). However, while the theoretical modelling is beginning to develop in this area, the transition to empirical work seems still difficult and the kind of econometric evidence usually cited in favour of *RBC* models is based mainly on theoretical work of the King-Plosser and Eichenbaum-Singleton variety.[1]

In what follows, three different kinds of empirical evidence offered in support of real and against monetary models of the cycle are briefly discussed. Firstly, recent empirical work on the trend/cycle decomposition of real output has emphasized the magnitude of (permanent) trend fluctuations, usually attributed to real shocks, at the expenses of (transitory) cyclical fluctuations, possibly due to monetary, or, more generally, nominal disturbances. Secondly, there is evidence on the *reverse causation* hypothesis on the money-output correlation based on the King-Plosser work. Finally, results from Granger-causality

[1] An interesting comparison of two alternative explanations of the correlation between real and monetary variables -a *RBC* model with endogenous money and a *credit-shock* model focusing on financial market imperfections- is carried out by Bernanke (1986). However, his results are not conclusively in favour of one theory as opposed to the other and appear to be sensitive to the detrending procedure adopted.

tests and the analysis of Vector Autoregressive (*VAR*) systems are often reported as favouring the *RBC* view.

1. Evidence on the non-stationarity of real variables. Starting from the seminal study by Nelson and Plosser (1982), a large amount of empirical work has addressed the issue of the non-stationarity of output and other real variables, and of the correct decomposition of output movements in a (stochastic) trend and a cyclical component. An earlier set of studies, focusing on *univariate* analyses of U.S. GDP, provided evidence of fluctuations in the *trend* component of output far larger than fluctuations in the *cyclical* component (Clark (1987), Campbell and Mankiw (1987), Cochrane (1988)). The presumption that monetary disturbances can determine only cyclical (transitory) fluctuations, whereas real shocks are the main determinant of (permanent) trend fluctuations, then led to the conclusion that monetary factors play a minor role in explaining output variability.

More recently, several objections to this view have been put forward at the statistical as well as theoretical levels. From a purely statistical perspective, Mc Callum (1986, 1989) and Christiano and Eichenbaum (1990) have emphasized the lack of power of unit root tests in discriminating between difference-stationary and trend-stationary series with a root close to, but less than, unity. Moreover, it has been shown that the persistence of shocks in univariate analyses cannot lead to conclusive evidence on the nature of the disturbances generating aggregate fluctuations (Cochrane (1990, 1991), Quah (1992)).

On the theoretical side, various authors have shown that it is possible to construct conventional monetary models of the cycle generating a highly persistent process for real output that could be erroneously attributed to the presence of real shocks. In fact, West (1988), using a variant of Taylor's (1980) staggered wage contracts model with shocks to monetary policy as the *only* source of instability, shows that a strong persistence of output is generated for plausible values of the model's basic parameters, when monetary authorities adopt a nominal interest rate rule. This result is extended by Phaneuf (1990) to a model of overlapping contracts with real wage objectives under fairly simple and general monetary policy processes. Even stronger results, in terms of output persistence, are obtained by Stadler (1986, 1990) in a purely monetary model (i.e. containing no exogenous technology shock) in which technological progress is endogenous, accumulated technical knowledge being positively related to past levels of output and employment. In this context it is shown that output may contain a root larger than unity and purely monetary innovations have a permanent effect on real activity.

2. Reverse causation evidence. Focusing on the role of the banking sector as producer of a particular type of intermediate good -monetary services- used as an input by all other sectors in the economy, King and Plosser (1984) have extended the original *n*-sector pure *RBC* model of Long and Plosser (1983) to include the financial sector as an additional industry in the economy. The basic mechanism which is responsible for the propagation across sectors of the exogenous technological shocks in the Long-Plosser economy (input-output interrelationships) is preserved and generates correlation between real and financial variables. This is due to a *reverse causation* effect: shocks originating in the real sector of the economy are transmitted to the financial sector mainly through the use of transaction services as an input in the production of final goods. What distinguishes this theory from earlier analyses of the endogeneity of money is a shift of the emphasis from the monetary authorities response to developments in the economy (as, for example, in Tobin (1970)) to the role of the private banking system.

The main empirical finding offered by King and Plosser in support of the reverse causation hypothesis for the U.S. over the 1953-1978 period, is based on contemporaneous regressions of the yearly rate of growth of output on the rate of growth of real and nominal monetary aggregates. Real activity appears to be much more correlated with inside money (real deposits) than with outside money (nominal monetary base). This evidence is broadly confirmed by Plosser (1991) on quarterly data over the extended 1948-1988 period: monetary aggregates such as M1 or M2 exhibit some correlation with real output whereas the monetary base appears consistently unrelated to real economic activity. However, in contrast with an *RBC* view, Lacker (1990) suggests that the observed correlation between inside money innovations and subsequent output movements may reflect anticipations of future (effective) monetary policy instead of being due to the production relationships between the banking sector and the other industries in the economy, with monetary services produced more rapidly than final goods to justify money movements leading output fluctuations.

3. Granger-causality and VAR evidence. The estimation and simulation of small *VAR* systems have now become standard techniques in empirical studies on the role of monetary factors in generating and shaping business cycle fluctuations (Bernanke (1986), Eichenbaum and Singleton (1986), Plosser (1991), Sims (1992), Eichenbaum (1992), among others).

Recent attempts to discriminate empirically between monetary and *RBC* theories of the cycle have also brought about the resurgence of Granger-causality tests as one of the main tools of analysis. Such tests had been at the centre of the debate on the effectiveness of demand management policies in the late 1970s until Buiter's (1984) demonstration that, in the cases of an optimizing controller and of *ad hoc* optimal feedback rules, policy instruments do not Granger-cause real endogenous variables, even though a change in the policy rule would affect their joint distribution function. This argument, implying that Granger-causality is *not necessary* for policy effectiveness, together with the parallel proof of *non-sufficiency* already offered by Sargent (1976)[2], led to the conclusion that no inference about policy effectiveness can be drawn from the results of such tests. In the new context of the debate over real *versus* monetary theories of economic fluctuations, the use of Granger-causality tests may appear, as noted by McCallum (1986, p.402), "*potentially appropriate*", since *RBC* theorists claim that not only anticipated movements in nominal variables but also the innovations in their processes are of no consequence for the behaviour of real activity. The proposition which seems to emerge from *RBC* models is therefore that output, employment, and other real variables will be block-exogenous to all nominal variables (prices, interest rates, money).[3] A finding of non-Granger-causality from nominal to real variables may thus appear to provide support for an *RBC* model of the economy.[4]

Additional evidence usually interpreted as favourable to a real view of fluctuations comes from innovation accounting exercises in small-scale *VAR* systems showing that only a relatively unimportant fraction of the variability of output and other real variables may be attributed to innovations in monetary aggregates. Eichenbaum and Singleton (1986) provided perhaps the most accurate and influential set of results from the application of *VAR* analyses (their findings are qualitatively confirmed by Litterman and Weiss (1985) and Boschen and Mills (1988)). Their underlying theoretical model is similar in nature to the basic *RBC* set-up, with preferences and technology as driving forces of the cycle, but money is introduced by means of a *cash-in-advance* constraint on consumers. The main testable implication is that money affects real activity through both its unanticipated and anticipated components.

Postwar U.S. monthly data are then examined to see whether they are consistent with such a model of the cycle or they favour a pure *RBC* view, attributing no real effect to

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[2] As shown by Sargent (1976), Granger-causality from money to output may be generated within a standard new classical model (in which only monetary *surprises* have real effects) through several channels: the influence of past monetary innovations in the semi-reduced form for output or the autocorrelation in real disturbances with only the contemporaneous money surprise entering the semi-reduced form for output.

[3] The issue (typical of early studies like Sims (1980b) and McCallum (1983b)) of whether nominal interest rates or some narrow monetary aggregate is the most appropriate indicator of monetary policy actions becomes immaterial in this new context.

[4] Conversely, evidence of nominal-to-real Granger-causality could be apparently interpreted against *RBC* theories, although not favouring *per se* any specific alternative model of fluctuations. This proposition has been challenged by Litterman and Weiss (1985) and King (1986), on the ground that nominal-to-real causality may be spurious -and hence not sufficient for rejection of *RBC* theories- if relevant variables have been incorrectly omitted from the empirical analysis.

movements in monetary aggregates. Two central findings are presented: *i)* money growth does not Granger-cause output growth, and *ii)* the percentage of the variance of output attributable to innovations in money supply growth is too small to justify the view that monetary policy shocks played a significant role as a driving force of the cycle in the U.S. over the postwar period. The overall message of Eichenbaum and Singleton is then that "*it would be difficult to construct a business cycle model which (a) assigns an important role to monetary factors, (b) is empirically plausible and (c) has the implication that money fails to Granger-cause output in the bivariate money-output relation*" (Christiano and Ljungqvist (1988), p.218).

In general, the finding of an insignificant contribution of money to the explanation of output fluctuations based on either Granger-causality tests or variance decomposition analyses is interpreted as evidence against monetary models of the cycle and in favour of real alternatives. The theoretical model developed in the next section is explicitly designed as a specific counterexample, showing how a *monetary* model of the Lucasian variety, extended to allow for a powerful stabilization role for monetary policy, may reproduce the main findings of the Eichenbaum-Singleton study.

The model presented in the next section is not aimed at yielding implications for all three kinds of empirical evidence briefly described above. Indeed, it specifically address the issue of the interpretation of money-to-output Granger-causality tests and output variance decomposition results in the context of the debate between "real" and "monetary" models of the cycle.

3. *Monetary stabilization policy in a "monetary" model of the cycle.*

The model presented in this section is an extension of the well-known Lucas (1973) *island* model, where we introduce expected inflation as a determinant of aggregate demand and thereby allow for a powerful role of systematic monetary policy, conducted on the basis of purely feedback rules, in stabilizing output fluctuations. Although supply shocks are present, the emphasis on aggregate demand (in particular monetary) disturbances makes the Lucas paradigm a typical "monetary" equilibrium model of the cycle, to be usefully contrasted with *RBC* alternatives.

3.1. *The structure and solution of the model.*

Following Lucas (1973) we consider an economy composed of a large number of separated, competitive markets indexed by $z = 1,\ldots,Z$, where a unique, homogeneous good is produced and demanded according to *local* supply and demand functions of the form (all variables are expressed in logarithms):

$$y_t^s(z) = \alpha \left(p_t(z) - E_z p_t \right) + \delta y_{t-1}(z) + \varepsilon_t \tag{1}$$

$$y_t^d(z) = m_t(z) - p_t(z) + \beta\left(\mathrm{E}_z p_{t+1} - p_t(z)\right) \qquad \beta > 0 \tag{2}$$

Local supply $y^s(z)$ depends positively ($\alpha > 0$) on the discrepancy between the realized local price $p(z)$ and the expectation of the economy-wide price level p formed by agents in the market. $E_z p_t$ denotes the rational expectation of p based on information available to agents in local markets at time t, including the structure of the model, past realizations of all variables and the contemporaneous local price $p_t(z)$; ϵ is an economy-wide white noise supply shock with mean zero and variance σ^2_ϵ. Local output supply depends positively ($0 < \delta < 1$) on lagged local output due to technological factors (e.g. adjustment costs and capital stock dynamics) of the kind emphasized by *RBC* theorists. This term captures, in an admittedly *ad hoc* way, persistence in output fluctuations. To justify (1), workers and firms may be assumed to observe directly only the price of their product and infer from this signal whether a change in their price reflects a change in the aggregate price level or indicates a change in relative prices. Only in the latter case workers will alter their labour supply and firms will adjust current production.[5]

Aggregate demand in each market is assumed to depend on the local real money supply $m(z)$ with unitary elasticity and on the locally expected inflation rate. Equation (2) can be interpreted as the reduced form of a standard *IS-LM* model where the real interest rate is a determinant of the *IS* curve whereas the nominal interest rate affects the LM curve and bond markets clear locally. It could also be interpreted as an inverted portfolio-balance equation à-la-Cagan.

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[5] As an alternative rationalization of (1), developing the insight of Friedman (1968), one can assume competitive local labour markets, where the demand for labour from profit maximizing firms with loglinear production functions is determined by the observed local producer real wage (local nominal wage deflated by the local price $p(z)$). Labour supply is an increasing function of the consumer real wage, i.e. the local nominal wage deflated by a consumer price index, not directly observable but inferred from all available information: $E_z p$. These assumptions on the labour market yield exactly a supply function of the form in (1). However, Bull and Frydman (1983) have pointed out some conceptual difficulties in integrating Friedman's discussion of the informational differences between employers and workers within the island paradigm with rational expectations.

The local nominal money supply is a stochastic fraction of the total money supply m:

$$m_t(z) = m_t + u_t(z) \tag{3}$$

where $u^d(z)$ denotes a white noise market-specific monetary shock with variance σ^2_z and the property that $\Sigma_z u^d(z) = 0$. Finally, total money supply is generated by the following feedback rule:

$$m_t = m_{t-1} + \mu v_{t-1} + v_t \tag{4}$$

where v denotes a white noise disturbance to money supply, independent of ϵ and $u(z)$ and with variance σ^2_v. The parameter μ captures the stabilizing response of monetary policy to past aggregate (monetary) shocks and will be optimally chosen so as to minimize output fluctuations around the full information level (to be made precise below).[6]

The model (1)-(4) is solved by means of the standard *undetermined coefficients* procedure (McCallum (1983a)). Substituting (3) and (4) into (2) and equating local demand and supply, we obtain the equilibrium local price level $p(z)$:

$$p_t(z) = \frac{1}{k}\left[\alpha \mathrm{E}_z p_t + \beta \mathrm{E}_z p_{t+1} + m_{t-1} + \mu v_{t-1} + v_t - \varepsilon_t + u_t(z) - \delta y_{t-1}(z)\right] \tag{5}$$

where $k = (1 + \alpha + \beta)$. In order to solve for the expectations of the aggregate price level in (5) we "guess" a solution for $p_t(z)$ of the following form:

$$p_t(z) = \pi_0 + \pi_1 u_t(z) + \pi_2 v_t + \pi_3 \varepsilon_t + \pi_4 v_{t-1} + \pi_5 y_{t-1}(z) + \pi_6 y_{t-1} \tag{6}$$

The local price level is assumed to depend on the whole set of aggregate and local contemporaneous disturbances, on the lagged economy-wide demand shock (via the monetary feedback rule) and on both the local and the aggregate (average) lagged output ($y \equiv (1/Z)\Sigma_z y(z)$). To form expectations of the aggregate demand and supply shocks, agents in each market are faced with a signal extraction problem. Given their information set, which includes $p_t(z)$, from (6) they can isolate the part of local price movements due to the composite contemporaneous disturbance $(\pi_1 u_t(z) + \pi_2 v_t + \pi_3 \epsilon_t)$, but cannot observe directly

[6] In (2) and (3) the only local and aggregate demand disturbances are the innovation in the money supply rule (4) and the local money shock $u(z)$. The introduction of additional (non-monetary) aggregate and local demand disturbances would merely complicate the algebra without altering our analysis and results. The same qualitative implications can be derived if monetary authorities are assumed to react to this demand shock or to past *supply* disturbances (ϵ_{t-1}).

each component of it. Knowing the stochastic distribution of the various shocks, the signal extraction problem can then be solved as:

$$\mathrm{E}_z v_t = \frac{\theta_1}{\pi_2}\left(\pi_1 u_t(z) + \pi_2 v_t + \pi_3 \varepsilon_t\right) \quad , \quad \mathrm{E}_z \varepsilon_t = \frac{\theta_2}{\pi_3}\left(\pi_1 u_t(z) + \pi_2 v_t + \pi_3 \varepsilon_t\right) \tag{7}$$

where

$$\theta_1 = \frac{\pi_2^2 \sigma_v^2}{\pi_1^2 \sigma_z^2 + \pi_2^2 \sigma_v^2 + \pi_3^2 \sigma_\varepsilon^2} \quad , \quad \theta_2 = \frac{\pi_3^2 \sigma_\varepsilon^2}{\pi_1^2 \sigma_z^2 + \pi_2^2 \sigma_v^2 + \pi_3^2 \sigma_\varepsilon^2} \tag{8}$$

Using (7) and (8), the final reduced form solution for the *aggregate* price level p is found to be (see *Appendix 1* for the complete derivation):

$$p_t = m_{t-1} + \pi(\mu)\cdot\left(v_t - \varepsilon_t\right) + \frac{\mu}{1+\beta}\, v_{t-1} - \frac{\delta}{1+\beta(1-\delta)}\, y_{t-1} \tag{9}$$

where $\pi(\mu)$ highlights the dependence of the coefficient on the contemporaneous composite aggregate shock (v_t-ϵ_t) on the policy parameter μ and is given by the following expression:

$$\pi(\mu) = \frac{\dfrac{\beta\theta_1}{1+\beta}\,\mu + \dfrac{\beta\delta\theta_2}{1+\beta(1-\delta)} + 1}{k - \alpha(\theta_1 + \theta_2)\left[1 - \dfrac{\beta\delta}{1+\beta(1-\delta)}(1-\theta_1-\theta_2)\right]} \tag{10}$$

Using (10) and (6)-(8) it is possible to derive the aggregate level of output (details in *Appendix 1*):

$$y_t = \alpha\pi(\mu)\cdot(1-\theta_1-\theta_2)\left(v_t - \varepsilon_t\right) + \delta y_{t-1} + \varepsilon_t \tag{11}$$

Aggregate output is determined by the monetary innovation v_t, the supply disturbance ϵ_t and lagged output. Due to the informational assumption of the model, the impact effect of aggregate disturbances depends on the variances of local and aggregate monetary and real shocks (σ^2_z, σ^2_v, σ^2_ϵ) through $\pi(\mu)$, θ_1 and θ_2. Our main focus is on the dependence of the impact coefficient in (11) on the policy parameter μ. In the following we analyze the role for stabilization policy in this extended Lucas framework and show the implications of different policy regimes for the interpretation of some empirical evidence apparently in favour of *RBC* theories of the cycle.

3.2. *Optimal monetary policy: theory and empirical implications.*

The assumed objective of the monetary authorities is the minimization of fluctuations of actual output y about its full information level y^*, obtained allowing agents to know immediately the aggregate or local nature of price level movements. Eliminating the agents' information problem, y^* is determined as:

$$y_t^* = \delta y_{t-1} + \varepsilon_t \tag{12}$$

Monetary authorities aim at minimizing the following conditional variance:

$$\mathrm{E}_{t-1}(y_t - y_t^*)^2 = \left[\alpha\pi(\mu)\cdot(1-\theta_1-\theta_2)\right]^2(\sigma_v^2+\sigma_\varepsilon^2) \tag{13}$$

where E_{t-1} denotes the expectation formed on the basis of the information set available to the monetary authorities, containing only lagged aggregate information. Minimizing (13) with respect to μ, the optimal policy parameter μ^* is:

$$\mu^* = -\frac{(1+\beta)\left\{1+\beta\left[1-\delta(1-\theta_2)\right]\right\}}{\beta\theta_1\left[1+\beta(1-\delta)\right]} < 0 \tag{14}$$

When $\mu=\mu^*$ perfect output stabilization is obtained and actual output y follows its full information path in (12).

The effectiveness of feedback monetary rules in this model is due to the presence of expected inflation as a determinant of demand, together with a "signal extraction" informational problem on local markets. Agents rationally attribute observed changes in local prices to disturbances of various kinds, according to (7) and (8). These changes will also alter inflation expectations and, in forming these revised expectations, the feedback money rule will be taken into account. An optimal response of the money supply to past shocks (μ^*) can thus affect inflation expectations in such a way as to offset completely the impact effect on demand of nominal shocks. The inability of agents to observe directly current disturbances creates some uncertainty about next period's monetary response, enabling the monetary authorities to stabilize real output even when they react to only one aggregate past disturbance (v_{t-1}), and not to the full set of shocks hitting the economy. The magnitude of the optimal response μ^* clearly depends on the relative variances of the various disturbances. In particular, we have:

$$\frac{\partial \mu^*}{\partial \sigma_v^2} = \frac{(1+\beta)\left\{[1+\beta(1-\delta)]\sigma_z^2 + (1+\beta)\sigma_\varepsilon^2\right\}}{\beta[1+\beta(1-\delta)]\sigma_v^4} > 0 \tag{15}$$

$$\frac{\partial \mu^*}{\partial \sigma_\varepsilon^2} = -\frac{(1+\beta)^2}{\beta[1+\beta(1-\delta)]\sigma_v^2} < 0 \tag{16}$$

$$\frac{\partial \mu^*}{\partial \sigma_z^2} = -\frac{1+\beta}{\beta\sigma_v^2} < 0 \tag{17}$$

If the variance of v is relatively high, a substantial fraction of all movements in the local price level will be interpreted as signalling aggregate demand shocks, to which money supply will respond in the next period. Perfect stabilization can thus be achieved with a relatively small response to demand disturbances (recall the negative value of μ^* from (14)). On the other hand, a larger response to v_{t-1} is needed to stabilize output if movements in the local price level are interpreted as signalling mainly either aggregate supply shocks (causing no monetary response) or local disturbances.

Moreover, the optimal value of the policy parameter depends in general on the stochastic properties of the various shocks. For example, relaxing our assumption of serially uncorrelated aggregate supply and demand disturbances and adopting simple *AR(1)* processes with coefficients $0<\rho_v<1$ and $0<\rho_\varepsilon<1$ respectively, it can be shown that a high degree of persistence of the demand shock, to which monetary policy reacts, generates a smaller (absolute) value for the optimal policy parameter μ^*. In this case, an innovation in v_t will influence also future values of the demand disturbance, triggering a stabilizing policy response not only at $t+1$ but also in subsequent periods. This is recognized by the agents and allows monetary authorities to reach perfect output stabilization with a smaller response to past shocks. The degree of persistence of the supply disturbance, to which monetary policy does *not* react, has the opposite effect of increasing the (absolute) value of μ needed for perfect output stabilization.[7]

It is now possible to derive some implications of the extended Lucas model for the evidence based on *VAR* systems and often used in the debate on real *versus* monetary theories of fluctuations. In this respect, the explicit consideration of different policy regimes is crucial in order to understand how these results are sensitive to changes in the policymakers' efforts to stabilize the economy. Two results are important here. First, considering a bivariate system with real output y and money supply m, under the perfect stabilization policy rule ($\mu=\mu^*$), there is no Granger-causality from money to output. In fact we have: $E(y_t|y_{t-1},\ldots)=E(y_t|y_{t-1},\ldots m_{t-1},\ldots)=\delta y_{t-1}$, so that past values of the money supply have no additional predictive power when added to past output values. Therefore, the evidence of absence of money-to-output Granger-causality is not sufficient to support the validity of real theories of the cycle, since this result may well be derived -as in the above model- from a different underlying structure, where purely nominal disturbances have a role and stabilization policy is effective.[8] Since also the interpretation of the opposite finding of money Granger-causing output as evidence against *RBC* theories in unwarranted (Litterman and Weiss (1985), King (1986)), the above result leads to the conclusion that causality tests are of little help in discriminating between alternative theories of economic fluctuations.

Different degrees of monetary policy stabilization would substantially alter the results of variance decomposition exercises (Sims (1980a, 1980b, 1982)) conducted using small-scale *VAR* systems. Using the moving average representation of output behaviour implied by (11), we can compute the overall variance of output and decompose it into two parts, attributable to the nominal (monetary) shocks v's and to the aggregate supply shocks ϵ's. The

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[7] Given the following stable *AR(1)* process for the monetary: $v_t=\rho_v v_{t-1}+\xi_t$, we have:

$$\mu^{v} = \mu^{*} + \rho_{v}\,\frac{\big(1+\beta(1-\delta)\big)\sigma_z^2 + (1+\beta)\sigma_\eta^2}{\big(1+\beta(1-\delta)\big)\sigma_\xi^2}$$

where μ^* is the optimal value of the policy parameter in (14) above, derived under the no autocorrelation assumption, and σ_ξ^2 is the variance of ξ. The similar assumption of an *AR(1)* process for the aggregate supply shock: $\epsilon_t=\rho_\epsilon\epsilon_{t-1}+\eta_t$, yields:

$$\mu^{\varepsilon} = \mu^{*} - \frac{\rho_\varepsilon(1+\beta)}{1+\beta(1-\rho_\varepsilon)}\,\frac{\sigma_\eta^2}{\sigma_\xi^2}$$

where σ_η^2 denotes the variance of η.

[8] Given the assumptions of the model, also when a non-stabilizing money rule is followed by the authorities ($\mu=0$) there is absence of money-to output Granger-causality. This may be generated, for example, by assuming autocorrelated demand disturbances. Under perfect stabilization, of course, such autocorrelation would not change the result of no Granger-causality between money and output.

proportion of the asymptotic output variance attributable to monetary innovations (V_m) is:

$$V_m = \frac{1}{1 + \left[\dfrac{(1-\alpha\pi(\mu))\sigma_z^2 + \sigma_\varepsilon^2 + \sigma_v^2}{\alpha\pi(\mu)\sigma_z^2}\right]^2 \dfrac{\sigma_\varepsilon^2}{\sigma_v^2}} \tag{18}$$

V_m clearly depends on the magnitude of π, which can be affected by the monetary rule adopted. Under perfect stabilization $\pi(\mu^*)=0$ and $V_m=0$: monetary innovations have no role in explaining output variance. This result is usually interpreted as supporting *RBC* views of the cycle, denying any influence of monetary variables (both anticipated *and* unanticipated) on output dynamics. However, the same result has been derived here from a model where monetary shocks may affect activity and monetary policy is extremely effective in stabilizing output. On the contrary, if we assume that the monetary authorities stick to a fixed money rule of the form ($\mu=0$), output follows (11) with $\pi(0)>0$ and $V_m>0$. Monetary innovations now have a detectable positive impact effect on output and would be attributed some weight in the decomposition of the asymptotic output variance.

In summary, our analysis highlights one channel -the countercyclical role played by monetary policy- whereby purely nominal innovations may be empirically attributed a negligible effect on output behaviour even if the underlying structure of the economy allows for such an effect.

The issue of what inferences on the underlying structure of the economy can be drawn from such innovation accounting exercises has already been addressed in the literature. In particular, various authors have shown how the detection of a significant role for monetary innovations in explaining output variance may well be generated by completely real models of fluctuations. The already mentioned Litterman-Weiss model offers one explanation for such a result, based on the correlation between monetary innovations and a (real) determinant of output excluded from the empirical analysis but observed by agents.

Similar conclusions are reached by King and Trehan (1984) in a model displaying full neutrality of money. Here, money supply responds endogenously to aggregate state variables -aggregate technological disturbances- that are not directly observable by agents. Therefore, any random movement of the money supply is partially attributed to the unobserved aggregate variable, which determines output fluctuations. The result is that monetary innovations appear to be correlated with output movements, although the underlying structure of the model denies any role for money. Following this insight, Siegel (1985) has provided a signalling model in which a completely neutral money supply yields valuable information about the level of real economic activity and future real interest rates.

Monetary innovations are viewed as signals, conveying some new information about real activity and hence correlation with output measures is the inevitable outcome of the conditional estimates of rational agents. Again, contemporaneous correlation between money supply and real output is made compatible with *RBC* theories.

On the other hand, the opposite empirical result that innovations in money supply are not capable of explaining a significant part of the variance of output is often regarded as clear evidence of the scarce empirical plausibility of monetary models of fluctuations. As the simple model discussed above shows, even the absence of correlation between monetary innovations and output can be compatible with a traditional monetary model of the cycle. Therefore, the conclusion we can draw from the foregoing discussion is that not only Granger-causality tests, but also innovation accounting techniques present rather serious problems when used to discriminate among competing macroeconomic theories.

The consideration of the behaviour of real variables under different policy regimes seems therefore a potentially fruitful way of discriminating among competing "monetary" and "real" theoretical models of fluctuations, since only the latter imply that shifts in the policy regime should not have any noticeable effect on real variables dynamics. The comparison made above, between a perfectly stabilizing feedback rule and a fixed money rule, is admittedly extreme but useful to illustrate the main points of our analysis. In practice, even less dramatic changes in policy rules, such as changes in the degree of policy countercyclicality, may be exploited for this purpose.[9]

In the next section we take up this empirical suggestion, comparing the results obtained for the U.S. in two time periods, characterized by a different degree of stabilizing effort in conducting monetary policy.

[9] Cross-country studies could also serve similar purposes. It is, in principle, possible to verify whether or not the impact effect of monetary factors on real output is inversely related to the degree of countercyclicality of the monetary rules adopted in each country.

4. A comparison of the interwar and postwar U.S. experience.

The aim of the present section is to assess to what extent the main implications of the theoretical "monetary" model of the cycle discussed in the preceding chapter in the context of the debate between "real" and "monetary" theories of fluctuations are supported by the data for the United States over different sample periods.

In particular, we are interested in the comparison between periods characterized by different attitudes of monetary authorities towards stabilization of the economy. According to the theoretical model, absence of money-to-output causality could be due to a successful stabilization policy. Therefore, in periods when policy has been more actively stabilizing, causality should be more difficult to detect than in other periods. Furthermore, the percentage of output variance attributable to monetary innovations should decrease when stabilization policy is more actively used.

To this aim, we identify the interwar (1922-1939) and the postwar (1952-1968) periods as characterized by different money rules followed by the authorities, with monetary policy systematically used to stabilization purposes only in the postwar period. Then, the time series properties of the data are investigated, with particular attention to the interwar period. Results from causality tests and output variance decomposition analyses are then reported for the two periods, and the differences in the real effect of monetary disturbances assessed.

Overall, our results do not contradict the view that activist monetary policy can have a sizeable effect in reducing the impact of nominal shocks on output, and may be responsible for some of the empirical evidence usually interpreted as favouring an *RBC* view of business cycle fluctuations.

4.1. The choice of the time periods and the characterization of their statistical properties.

In the recent literature, several authors have provided extended accounts of the macroeconomic performance of the U.S. economy over different historical periods. One aspect which is often debated is the extent of and the explanation for the apparent reduction in the severity of business cycle fluctuations experienced after World War II. Until recently, the belief in a substantial reduction of the average amplitude of business cycles in the postwar period was widely shared among economists. Both Taylor (1986) and DeLong and Summers (1986), comparing the pre-1914 with the post-1945 periods, detect -during the latter- a remarkable improvement in the macroeconomic performance of the U.S. economy, especially in terms of a decrease of the cyclical variability of output. However, they offer different explanations for this finding. Taylor attributes such improvement mainly to a reduction of the impulses generating cyclical fluctuations, notwithstanding the increased rigidity of prices and wages in the postwar period, whereas DeLong and Summers conclude that the postwar economy displayed smaller fluctuations because of greater public and private effort to smooth consumption and an increased degree of price rigidity (in turn due to the increased institutionalization of the economy).

A different position, based on a radical reconstruction of U.S. historical data on unemployment, GNP and industrial production, has recently been taken by C. Romer (1986, 1989). She argues that the alleged volatility of the economy in the earlier periods is overstated, mainly because of the particular assumptions underlying the construction of official macroeconomic series.[10] Finally, in their thorough analysis of U.S. business cycles in historical perspective, DeLong and Summers (1988), using C. Romer's prewar data and including the interwar years, present evidence of a considerable improvement in macroeconomic performance in the post-World War II period, and attribute this improvement to successful stabilization policies.

Notwithstanding the variety of opinions on the effectiveness of stabilization policy, there is little doubt that demand management was actively employed only in the period following the second World War. It seems therefore useful, for our purposes, to compare particular aspects of the macroeconomic performance of the U.S. economy over the interwar and the postwar periods. Our attention is focused on monetary policy, which underwent a gradual evolution since the founding of the Federal Reserve System in 1914. During the 1920s and 1930s Federal Reserve decision makers "*gradually came to understand what effects the system's open market purchases and sales of government securities had in the new world of fractional reserve banking directly based on central bank liabilities.*" (B. Friedman (1986), p.399). In 1923 the body which will evolve into the modern Federal Open Market Committee was created, leading temporarily to an "*increasing emphasis on open market operations in a monetary policy context, but in the 1930s the confusion of the depression and the associated international monetary crisis, including the abandonment of the gold standard in 1934, arrested the developments of the monetary policy mechanism*" (p.400). On the whole, monetary policy was not aimed at output stabilization purposes over the interwar period and the lack of reaction of monetary authorities in the face of the developments in the

[10] Sheffrin (1989, ch.2) provides a thorough critical discussion of C. Romer's contributions. Balke and Gordon (1989), applying a different methodology, construct an alternative series for prewar GNP, which displays as much variability as the traditional series.

real economy at the onset of the Great Depression are often viewed as a major evidence of this behaviour (Friedman and Schwartz (1963), C. Romer and D. Romer (1989)).[11]

In the aftermath of the World War II until the Treasury-Federal Reserve Accord in 1951, U.S. monetary authorities maintained an obligation to support the open market price of the government's outstanding debt, and only after the 1951 Accord the Federal Reserve started to play an independent macroeconomic role, actively reacting to the developments in the real economy in a stabilizing manner (Meulendyke (1988)). Over the whole postwar period, however, the monetary authorities also tried to control inflation. Whereas the latter objective was of secondary importance in the 1950s and 1960s, due to the low inflation level, in the 1970s and 1980s, it became monetary policy's central concern. C. Romer and D. Romer (1989), analyzing in detail the behaviour of U.S. monetary authorities over the postwar period with the final aim of testing for the real effects of monetary disturbances, identify several episodes in which monetary authorities attempted to exert a contractionary influence on the economy in response to excessively high inflation rates, reacting more than it would have been necessary in order to offset perceived or expected increases in aggregate demand. Five such episodes are identified since 1951: September 1955, December 1968, April 1974, August 1978, and October 1979. Moreover, from the evidence presented in the Romers' study, it seems that the extent of the real effects of the 1955 anti-inflationary monetary policy reaction is much smaller, at least in the two years following the shift, than that in any of the post-1968 similar episodes. Therefore, there is evidence that the response to high and rising rates of inflation became more intense from the end of 1968, if compared with the first part of the postwar period.

On these grounds, we adopt the view that monetary authorities have pursued a policy mainly aimed at output stabilization from the 1951 Accord until at least the end of 1968, whereas output stabilization was not a major concern of the monetary authorities in the interwar (1922-1940) period. In *Appendix 2* a quantitative evaluation of the different degree of countercyclicality of monetary policy in the two periods is conducted by estimating simple feedback rules for the growth rate of the M1 money stock, trying to capture the systematic response of monetary authorities to the observed state of the economy. The results lend some support to the view of a stronger stabilization effort in the 1952-1968 period.

Prior to proceed with the analysis of small-scale *VAR* systems, we investigate the

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[11] A considerable debate has raged over whether the Federal Reserve failure to respond appropriately to the Great Depression was due to policymakers' incapability of understanding that more decisive intervention was necessary, or was caused by a deliberately chosen policy. Wheelock (1991, 1992) provides an evaluation of the literature on this isssue.

time-series properties of the data used. In fact, results from variance decompositions and causality tests reported in several studies of U.S. postwar data appear dramatically sensitive to different detrending procedures and to whether levels instead of first differences of the series are used (Eichenbaum and Singleton (1986), Bernanke (1986), Christiano and Ljungqvist (1988)). For example, Eichenbaum and Singleton (1986, tables 5.2-5.3), using a three-variable system including industrial output, the price level and money supply over the 1959-1983 period, find that monetary innovations account for 11% of the 48-month forecast error variance of real output when first differences of the data are used, whereas 33% of output variance can be attributed to monetary innovations if the variables are linearly detrended. Given the extent of the problem, a preliminary characterization of the time trend and unit root properties of the data is in order. Moreover, recent developments in the study of non-stationary time series have shown that both the asymptotic and the finite-sample distributions of causality tests are sensitive to the presence of unit roots and time trends in the series (Sims, Stock and Watson (1990)). Building on such theoretical results, Stock and Watson (1989), Krol and Ohanian (1990) and Friedman and Kuttner (1993) have developed and applied a sequential testing procedure to characterize empirically the behaviour of money, output, prices and interest rates. This procedure, making use of augmented Dickey-Fuller (*ADF*) tests, is applied to our data, including nominal M1, m, real GNP, y, the GNP deflator p and the 4- to 6-month commercial paper nominal interest rate, R. All data are quarterly, seasonally adjusted, and, with the exception of R, expressed in logarithms. In what follows, only the main results of our analysis (reported in more detail in *Appendix 2*, together with data sources) are summarized, and some specific problems discussed.

The results of the univariate and multivariate tests for the *postwar* period show that all variables are stationary in first differences and only for the growth rate of money (Δm) there is evidence of a linear time trend. Therefore, the following specification of our series is adopted for the postwar period:

$$\begin{aligned}
\Delta y_t &= \alpha_{y0} + \eta_t \\
\Delta m_t &= \alpha_{m0} + \alpha_{m1} t + \psi_t \\
\Delta p_t &= \alpha_{p0} + \omega_t \\
\Delta R_t &= \rho_t
\end{aligned}$$

where $\Delta\eta$, $\Delta\psi$, $\Delta\omega$, and $\Delta\rho$ are mean zero stationary processes. Now, letting X denote the vector of variables included in each of the systems considered -(y,m), (y,m,p), and (y,m,p,R)- and assuming that the corresponding sub-vector of (η,ψ,ω,ρ), ξ, has a *VAR(n)* representation of the form:

$$A(L)\xi_t = \chi_t$$

where χ is a vector of innovations, then the systems may be written as:

$$A(L)\Delta X_t = \gamma_0 + \gamma_1 t + \chi_t \tag{19}$$

where $\gamma_0+\gamma_1 t=A(L)(\alpha_0+\alpha_1 t)$. This *VAR* representation of the variables is consistent with the trend and integration properties of the data and therefore provides a valid framework for the application of Granger-causality tests and variance decomposition techniques. In particular, the inclusion of a time trend in (19) seems especially important since, as shown by Stock and Watson (1989), failure to consider this trend tends to obscure the Granger-causal relationship between money and output in the United States.

For the *interwar* period, a statistical representation of the variables in first differences without deterministic terms seems appropriate, given the results of the unit root tests reported in *Appendix 2*. However, the interwar period poses some problems for the implementation of unit root tests, since it includes a sub-period -corresponding to the years of the Great Depression (1929-1933)- when output collapsed and also the time-series behaviour of other macro variables was altered. The *ADF* test procedure used in *Appendix 2* is implicitly based on the view that this episode, notwithstanding its magnitude, is part of the realization of the underlying process generating macroeconomic time series, and is not due to exogenous forces altering such process. On the contrary, if one adopts this alternative view, considering the Great Depression as the consequence of an exogenous change in the data generating process around 1929, the testing procedure employed to characterize the time series properties of the data has to be appropriately modified. In particular, the null hypothesis of difference-stationarity and the alternative hypothesis of trend-stationarity used in the formulation of the *ADF* unit root tests must be revised. As shown by Perron (1989), one could hardly reject the unit root hypothesis on the basis of such tests even if the series are stationary around a linear time trend, but with a one-time change in their level (the *crash* hypothesis) or in the trend coefficient (the *changing growth* hypothesis). Taking into account the possibility of both kinds of break in the deterministic trend of the series, the null hypothesis to be tested may be reformulated, for the generic variable z, as follows:

$$H_0: \quad z_t = a_1 + z_{t-1} + dD(T_B)_t + (a_2 - a_1)DU_t + e_t \tag{20}$$

where the inclusion of the dummy variable $D(T_B)_t$ (set equal to *1* for $t=T_B+1$ and to *0* for the rest of the sample) allows for an exogenous change in the level of the series at the break date T_B, measured by the coefficient d, whereas the variable DU (set to *0* for $t \leq T_B$ and to *1* for $t>T_B$) captures an exogenous change in the series growth rate, measured by a_2-a_1. The stochastic term e is assumed to be generated by a stationary *ARMA(p,q)* process. Under H_0, z follows a process with a unit root and a shift in the level and the growth rate at time T_B. Consequently, the alternative hypothesis of trend-stationarity is reformulated as:

$$H_1 : \quad z_t = a_1 + b_1 t + (a_2 - a_1) DU_t + (b_2 - b_1) DT_t^* + e_t \tag{21}$$

where DU is as defined above and $DT^* = t - T_B$ for $t > T_B$ and 0 for $t \leq T_B$. Under H_1, z is stationary around a linear time trend, with slope b_1 for $t \leq T_B$ and b_2 afterwards. Also at T_B, the level of the series has a shift measured by $a_2 - a_1$.

In order to allow for a gradual reaction of the economy to exogenous breaks in the trend function, the following specification of the time series process for z, nesting both hypotheses, is adopted for testing:[12]

$$\Delta z_t = a + \alpha z_{t-1} + \theta DU_t + bt + \gamma DT_t^* + dD(T_B)_t + \sum_{j=1}^{k} c_j \Delta z_{t-j} + e_t \tag{22}$$

The test of the null hypothesis that $\alpha = 0$ against the alternative ($\alpha < 0$) is then performed by comparing the computed t-statistic for α with the critical values provided by Perron (1989, table IV).

An important feature of this testing procedure is the assumed *a priori* knowledge of the date (T_B) of the potential structural changes in the process generating z. In practice, the observed behaviour of the series and other relevant information may suggest a precise dating for T_B. On the contrary, if one interprets apparent anomalies in the series as realizations from the "tails" of the distribution of the data generating process and not as exogenous events, the testing procedure has to allow for trend breaks as in (22) but occurred at dates unknown *a priori*. Zivot and Andrews (1992) have recently proposed a modification of Perron's test so as to endogenize the break date. The transformation of the test in (22), which is conditional on the choice of T_B and therefore data-dependent, into an unconditional test is obtained by reformulating the null hypothesis to be tested, eliminating from (20) the two dummy variables ($D(T_B)$ and DU) capturing structural breaks. The alternative hypothesis and the nesting equation retain the formulations in (21) and (22), but without assuming a known T_B. Equation (22) is then estimated for all possible dates T_B (only excluding short periods at the beginning and at the end of the sample) and the test is conducted on the lowest estimated α. In so doing, the break date most favourable to the (alternative) hypothesis of stationarity around a deterministic trend displaying structural breaks is selected. Critical

[12] The assumption that the economy displays the same response to shocks in the trend function and to any other shock is implicit in (22).

values for the test are provided by Zivot and Andrews (1992, Table 4A).

The Zivot-Andrews version of the testing procedure outlined above is applied to our data for the 1922-1940 period to assess the validity of the chosen *VAR* system representations. In Table 1 we report, for each variable, the two lowest values of the t-statistic on the coefficient α (t_α) in (22) and the corresponding date for T_B. All series show low values of t_α in some quarter from the end of 1930 to the beginning of 1932, but in no case these values are statistically significant at the 5% level.[13] [14] Therefore the results of the unit root tests motivating our choice of a *VAR* system in first differences with no additional trend term, seems robust to the above alternative assumptions on the nature of the deterministic trend.

Table 1
Zivot-Andrews test: 1922-1940

Variable	*Lowest values of t_α and corresponding dates:*				*Q(8)*
	1		*2*		
	t_α	*Date*	t_α	*Date*	
y	-3.65	1931.2	-3.25	1931.3	4.8
m	-4.11	1931.1	-3.90	1930.4	2.9
p	-5.02	1930.4	-5.01	1931.3	10.8
R	-4.65	1931.4	-4.19	1932.1	13.6

Note: The 5% critical value of the Zivot-Andrews test described in the text is -5.08. *Q(8)* denotes the Lijung-Box test for eighth-order serial correlation in the residuals of equation (22), distributed as $\chi^2(8)$ (5% critical value: 21.0). Critical values of the Perron (1989) test are: -4.24 (5%) and -4.89 (1%).

[13] Even considering the critical values tabulated in Perron (1989) for a known break date (chosen to correspond to the quarter with the lowest t_α), only for the price level p is the null hypothesis of stochastic trend rejected at the 1% level.

[14] Using yearly data over a much longer time span, Perron (1989), Zivot and Andrews (1992) and Ben-David and Papell (1994) found evidence of a trend break in 1929 for the GDP series.

4.2. VAR analysis of the two periods: results.

Having established a satisfactory time-series representation of the variables, we can now implement causality tests and variance decomposition analyses on *VAR* systems including first only output and money, and then extended to the price level and the interest rate, over the 1922-1940 and 1952-1968 sample periods. Similar analyses have been conducted in the literature, among others, by Sims (1980b, 1982) and, for the postwar period only, by Eichenbaum and Singleton (1986).

A first set of results is displayed in Table 2, panel A, where the *F*-statistics (with the associated significance levels) from testing the null hypothesis that the rate of growth of money fails to Granger-cause output growth are reported. Although the theoretical "monetary" model of the cycle discussed in the preceding section does not yield predictions about causality tests in terms of the level of statistical significance of the estimated parameters, a finding of a greater degree of money-to-output Granger-causality in the postwar period (with an active monetary policy) with respect to the interwar years would cast serious doubts on the applicability of such a model to our data. Therefore, we interpret the results from causality tests as a broad check on the admissibility of the monetary model as a valid alternative to a *RBC* interpretation of the data. Strong evidence of Granger-causality from money growth to output growth is detected in the interwar period for all three system specifications, with significance levels of the *F*-statistics always below 5%. In the postwar period, the degree of money-to-output causality decreases sharply, with significance levels of the test ranging from 12 to 17%. Overall, the finding of a weaker evidence of causality in the 1952-1968 period does not contradict the view that a more active stabilization policy may be responsible for the absence of Granger-causality from money to output, as shown in the Lucas-type model of the previous section.

However, stronger implications are derived from that model with respect to output variance decompositions in *VAR* systems. Letting x_t be the vector of the variables in the *VAR* (Δy, Δp, Δm and ΔR in the four-variable case), the *structural* form of the system may be written as:

$$A x_t = B(L) x_{t-1} + \varepsilon_t \tag{23}$$

where matrix A (with ones on the diagonal) describes the contemporaneous relations among the variables, $B(L)$ is a matrix of polynomials in the lag operator and ϵ_t is the vector of structural disturbances, with $E(\epsilon_t \epsilon_t') = \Sigma_\epsilon$. The residual variance matrix Σ_ϵ is assumed diagonal, implying orthogonality of the structural disturbances. The estimated *VAR* system is the reduced form of (23), given by:

$$x_t = C(L)x_{t-1} + u_t \qquad C(L)=A^{-1}B(L) \;, \; u_t=A^{-1}\varepsilon_t \tag{24}$$

The *VAR* residuals in u are linear combinations of the underlying structural disturbances with the following non-diagonal covariance matrix: $E(u_t u_t')=\Sigma_u=A^{-1}\Sigma_\varepsilon A^{-1'}$. Therefore some structural assumptions are needed in order to decompose the vector of estimated reduced-form residuals into orthogonal components, to be interpreted as innovations to each variable in the system. Various orthogonalization procedures have been applied in the literature by Sims (1980a), Blanchard and Watson (1986), Bernanke (1986), Blanchard (1989) and Blanchard and Quah (1989). Here we adopt the method originally proposed by Sims (1980a) and employed also by Eichenbaum and Singleton (1986), based on a simple Choleski factorization of the *VAR* residual covariance matrix. The implied structural model has a recursive form, with the residual from equation i in the system expressed as a linear combination of residuals from equations $1,2,\dots,i-1$ only. The matrix A of contemporaneous relations is therefore assumed lower-triangular, with ones on the diagonal.[15] The ordering of variables then reflects beliefs on the underlying theoretical model of the economy. According to an *RBC* view of cyclical fluctuations, for example, real variables should appear before nominal variables in the orthogonalization: no contemporaneous impact of nominal on real quantities is allowed, whereas part of the innovations in nominal variables is attributed to real disturbances. In the following empirical analysis this preferred *RBC* ordering is adopted, with output growth being placed first in all specifications. When included in the system, inflation and the interest rate change are ordered second and fourth. Finally, the money growth rate is placed last in the bivariate and trivariate systems, and precedes only ΔR in the complete *VAR*. The contemporaneous correlation between money and output growth innovations is then given an output-to-money interpretation, consistent with potential money supply endogeneity, emphasized by *RBC* theorists as an explanation for the observed comovements of money and output. The evidence of a negative correlation between Δm and ΔR in both periods supports the view that money growth innovations reflect money supply disturbances (having a "liquidity" effect on interest rates), motivating the

[15] The different procedure implemented by Blanchard and Watson (1986), Bernanke (1986), and Blanchard (1989) uses a set of structural assumptions with a precise economic rationale to go from the reduced-form residuals to uncorrelated structural innovations, instead of adopting an implicitly lower-triangular structural model as in Sims (1980a). However, Eichenbaum and Singleton (1986) criticize this procedure on the ground that for a large class of dynamic models the parameters of the innovation covariance matrix cannot be identified separately from the parameters of the reduced-form equations. As a further alternative, Blanchard and Quah (1989) employ long-run restrictions on the dynamic responses of the endogenous variables to different innovations to identify structural disturbances.

choice of ordering money before interest rate in the complete system.[16]

Results from the forecast error variance decomposition for output growth are shown in Table 2, panel B, where the percentage of the output variance attributable to money growth disturbances is reported for various time horizons (1, 3 and 5 years), together with 70% confidence bounds. The comparison between the two periods considered shows that the fraction of output variance attributable to nominal (monetary) shocks is consistently lower in the postwar years, when monetary policy was actively used for stabilization purposes. The pattern of results is consistent across all system specifications. In the bivariate *VAR*, though the point estimates show a reduction from 20 to 10% in the contribution of money shocks, the relatively wide confidence intervals do not allow sharp inferences. The inclusion of the price level and the interest rate in the system makes the result statistically more reliable, with point estimates of 28 and 8% in the interwar and postwar periods respectively.

For the complete systems, Table 3 presents the estimated elements of the matrix A, capturing the contemporaneous relations among the variables. Only some coefficients are statistically significant, with a positive response of money to output only in the interwar years and a negative reaction of the interest to money growth in both periods.

In order to assess the robustness of the above results, two variants of the four-variable *VAR* system have been considered. First, we tried a different ordering of the variables, suggested by a particular interpretation of the relationships between money and income in the *RBC* spirit. As mentioned above, according to *RBC* theories, the money-output comovement is mainly due to the endogeneity of monetary aggregates, reacting to changes in production. This "reverse causation" argument might also explain the empirical finding of money leading output, besides the contemporaneous correlation between the two quantities. Our estimated *VAR*, with output ordered before money, attributes the contemporaneous correlation to an endogenous response of the monetary aggregate to output innovations. However, if in reality this endogenous reaction leads observed output movements, an ordering with also the interest rate preceding money could be more appropriate. In fact, interest rate innovations could reflect new information available to agents in financial markets about future output behaviour, in anticipation of which monetary aggregates may react. With the interest rate placed before money in the *VAR* such leading role is attributed to interest rates (whose innovations are not interpreted here necessarily as

[16] A negative correlation with interest rate innovations is viewed as a minimum requirement for interpreting innovations to money as money supply disturbances by Todd (1990), Sims (1992) and Eichenbaum (1992). Impulse response functions confirm that for both periods money disturbances generate a (temporary) negative reaction in interest rates.

purely nominal disturbances) and not to money supply movements. When the complete *VAR* is estimated with this alternative ordering the results in Table 2 are confirmed: the fraction of output variance attributable to nominal disturbances is 26% in the interwar period and only 3% in the postwar years.

A further check on the robustness of our results concerning the interwar period is suggested by some recent studies, starting from Bernanke (1983), emphasizing the role of banking crises in determining the depth of the Great Depression. Default by large banks, concentrated particularly at the beginning of the 1930s, and the consequent disruption of the payment and financial intermediation system have caused changes of the money stock that can hardly be considered of a purely nominal nature. From this perspective, the results obtained above may overestimate the importance of nominal monetary disturbances in the interwar period, attributing to innovations in Δm also the real effects of banking crises. To assess this possibility, we extended the system to include a variable capturing the extent of the banking crises: the real value of deposits in suspended banks over the 1922-1940 period. As shown in Figure 1, this variable increases sharply at the beginning of the 1930s, when money supply displays a marked decline.[17] This series (in log differences) is ordered before money in the *VAR* so as to emphasize its role in the explanation of output variance with respect to money. The results show that, although some 12% (±5.5%) of output variance can be attributed to shocks to the financial intermediation and payment system, the fraction due to monetary disturbances is almost unchanged at 26.5% (±7.2%).

Overall, our main results seem robust also to this extension of the estimated system and do not contradict the message of the theoretical model of section 3, that an effective (monetary) stabilization policy may be responsible for some empirical findings apparently in favour of *RBC* theories.

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[17] This variable, obtained from the *Federal reserve Bulletin*, is used also by McCallum (1990) and Bordo, Choudhri and Schwartz (1993) in a different context.

Table 2
VAR analysis.

A. *Money-income causality tests: F-statistics (p-values).*

Sample period:	*System:*		
	Δy, Δm	Δy, Δm, Δp	Δy, Δm, Δp, ΔR
1922(1)-1940(4)	2.91 (0.02)	4.47 (0.01)	4.66 (0.001)
1952(1)-1968(4)	1.71 (0.16)	1.94 (0.12)	1.67 (0.17)

Note: The *F*-statistics test the hypothesis that the coefficients on four lags of Δm are jointly zero in the output equation corresponding to the three systems considered and containing four lags of the included variables, a constant and, for the 1952-1968 period only, a linear time trend.

B. *Forecast error variance decomposition for* Δy*: the role of money growth innovations.*

(i) *Period: 1922-1940*

Forecast horizon (quarters):	*System:*		
	Δy, Δm	Δy, Δp, Δm	Δy, Δp, Δm, ΔR
4	17.3 ($\pm$ 7.2)	21.2 ($\pm$ 7.7)	21.7 ($\pm$ 7.6)
12	20.0 ($\pm$ 8.6)	26.2 ($\pm$ 8.8)	28.3 ($\pm$ 8.8)
20	20.0 ($\pm$ 9.0)	26.2 ($\pm$ 8.9)	28.3 ($\pm$ 8.9)

(ii) *Period: 1952-1968*

Forecast horizon (quarters):	*System:*		
	Δy, Δm	Δy, Δp, Δm	Δy, Δp, Δm, ΔR
4	8.9 ($\pm$ 6.7)	9.3 ($\pm$ 6.6)	6.9 ($\pm$ 5.8)
12	9.9 ($\pm$ 7.5)	10.3 ($\pm$ 7.5)	7.6 ($\pm$ 6.5)
20	9.9 ($\pm$ 7.6)	10.4 ($\pm$ 7.6)	7.8 ($\pm$ 6.7)

Note: Each entry shows the percentage of the forecast error variance of Δy attributable to Δm at various forecast horizons. The ordering used in the orthogonalization of the residual matrix is indicated in the second row. System specifications maintain the statistical characterization of the data discussed in section 4.1. 70% confidence bounds are reported in parentheses.

Table 3
Estimated contemporaneous relations in the four-variable VAR systems.

(i) Period: 1922-1940

$$u_{yt} = \varepsilon_{yt}$$

$$u_{pt} = \underset{(0.049)}{0.194}\, u_{yt} + \varepsilon_{pt}$$

$$u_{mt} = \underset{(0.072)}{0.227}\, u_{yt} - \underset{(0.155)}{0.009}\, u_{pt} + \varepsilon_{mt}$$

$$u_{Rt} = \underset{(0.016)}{0.029}\, u_{yt} + \underset{(0.033)}{0.006}\, u_{pt} - \underset{(0.025)}{0.099}\, u_{mt} + \varepsilon_{Rt}$$

$$\sigma_y = 0.025 \qquad \sigma_p = 0.0104 \qquad \sigma_m = 0.014 \qquad \sigma_R = 0.0029$$

(ii) Period: 1952-1968

$$u_{yt} = \varepsilon_{yt}$$

$$u_{pt} = \underset{(0.056)}{0.026}\, u_{yt} + \varepsilon_{pt}$$

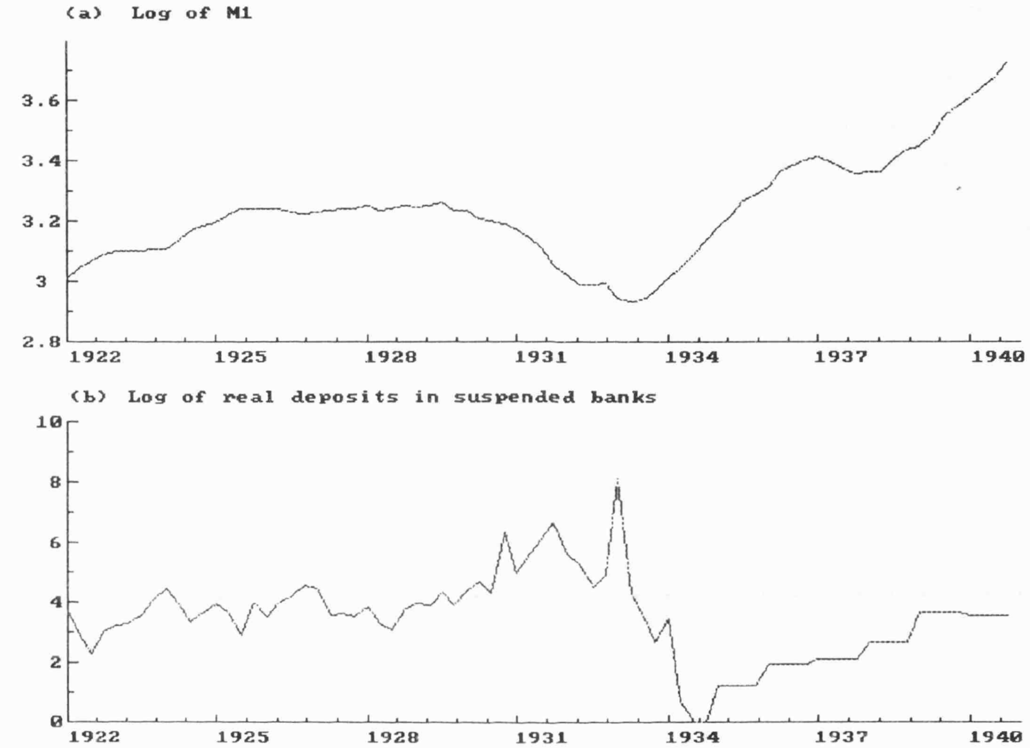
$$u_{mt} = \underset{(0.052)}{0.067}\, u_{yt} - \underset{(0.119)}{0.001}\, u_{pt} + \varepsilon_{mt}$$

$$u_{Rt} = \underset{(0.033)}{0.154}\, u_{yt} + \underset{(0.074)}{0.136}\, u_{pt} - \underset{(0.077)}{0.219}\, u_{mt} + \varepsilon_{Rt}$$

$$\sigma_y = 0.0071 \qquad \sigma_p = 0.0031 \qquad \sigma_m = 0.0031 \qquad \sigma_R = 0.0019$$

Note: The table reports the estimated coefficients of the lower-triangular A matrix in the following relation: $Au_t = \epsilon_t$, where u and ϵ are the vectors of *VAR* residuals and of the orthogonalized disturbances respectively (see equation (24) in the text). σ denote the standard error of the structural-form disturbances in ϵ.

Figure 1
Money supply and deposits in suspended banks (1922-1940).



5. Conclusions.

This chapter has addressed the issue of what inferences concerning the underlying structural model of the economy may be derived from the results of causality tests and variance decomposition exercises widely applied in the empirical macroeconomic literature. Advocates of *RBC* theories of fluctuations are inclined to interpret the absence of Granger-causality from nominal (in particular, monetary) variables to real quantities and, more important, the finding that only a small proportion of output variance may be attributed to monetary innovations in the analysis of *VAR* systems, as evidence against monetary models of the cycle. The extended Lucas model analyzed in this chapter provides an example of a "monetary" equilibrium model of the cycle which is capable of generating the kind of evidence usually interpreted as favouring the *RBC* view. Absence of money-to-output causality and of correlation between monetary innovations and output are here the by-products of a successful stabilization policy.

The simple empirical implication is that the analysis of periods characterized by widely alternative policy regimes, with different stabilizing stance, should detect a greater role for nominal disturbances in determining output variability when monetary policy is more actively (and effectively) used for stabilization purposes. The results obtained from the comparison of the interwar and the postwar (until 1968) periods for the U.S., with monetary policy being employed to pursue output stabilization only in the latter, show a substantial reduction in the proportion of output variance accounted by monetary innovations in the postwar years.

Even though this pattern of results may be consistent with "monetary" business cycle models other than the one we adopted, it seems difficult to convincingly account for it in the framework of the *RBC* theory of fluctuations.

Appendix 1

In this appendix we derive the solution for the aggregate price level p and output y given in equation (9) and (11) in section 3 (equations identical to those in the text maintain their original numbering).

From the guess solution for $p_t(z)$ given in (6), aggregating over all markets and recalling that $(1/Z)\Sigma_z u(z)=0$ and $y \equiv (1/Z)\Sigma_z y(z)$, we obtain:

$$p_t = \frac{1}{Z}\sum\nolimits_z p_t(z) = \pi_0 + \pi_2 v_t + \pi_3 \varepsilon_t + \pi_4 v_{t-1} + (\pi_5 + \pi_6) y_{t-1} \qquad \text{(A1)}$$

Taking conditional expectations of (A1):

$$\mathrm{E}_z p_t = \pi_0 + \pi_2 \mathrm{E}_z v_t + \pi_3 \mathrm{E}_z \varepsilon_t + \pi_4 v_{t-1} + (\pi_5 + \pi_6) y_{t-1} \qquad \text{(A2)}$$

Making use of the solution to the signal extraction problem given in (7) and (8) and substituting fot $E_z v_t$ and $E_z \epsilon_t$, (A2) becomes:

$$\mathrm{E}_z p_t = \pi_0 + (\theta_1 + \theta_2)\left(\pi_1 u_t(z) + \pi_2 v_t + \pi_3 \varepsilon_t\right) + \pi_4 v_{t-1} + (\pi_5 + \pi_6) y_{t-1} \qquad \text{(A3)}$$

Leading (A1) by one period we have:

$$p_{t+1} = \pi_0 + \pi_2 v_{t+1} + \pi_3 \varepsilon_{t+1} + \pi_4 v_t + (\pi_5 + \pi_6) y_t \qquad \text{(A4)}$$

In order to derive $E_z p_{t+1}$, we use the fact that, given the assumed white noise properties of v and ϵ, $E_z v_{t+1} = E_z \epsilon_{t+1} = 0$. Then, we only have to compute $E_z y_t$. To this aim, we aggregate the supply function (1) over all markets, obtaining:

$$y_t = \alpha\left[p_t - \frac{1}{Z}\sum\nolimits_z \mathrm{E}_z p_t\right] + \varepsilon_t + \delta y_{t-1} \qquad \text{(A5)}$$

Using (A3) and averaging across markets we get:

$$\frac{1}{Z}\sum\nolimits_z \mathrm{E}_z p_t = \pi_0 + (\theta_1 + \theta_2)\left(\pi_2 v_t + \pi_3 \varepsilon_t\right) + \pi_4 v_{t-1} + (\pi_5 + \pi_6) y_{t-1} \qquad \text{(A6)}$$

Subtracting (A6) from p_t in (A1) yields:

$$p_t - \frac{1}{Z}\sum\nolimits_z \mathrm{E}_z p_t = [1 - (\theta_1 + \theta_2)]\left(\pi_2 v_t + \pi_3 \varepsilon_t\right) \qquad \text{(A7)}$$

Finally, substituting (A7) into (A5) we can derive the aggregate output equation as:

$$y_t = \alpha[1 - (\theta_1 + \theta_2)]\left(\pi_2 v_t + \pi_3 \varepsilon_t\right) + \varepsilon_t + \delta y_{t-1} \qquad \text{(A8)}$$

The conditional expectation $E_z y_t$ is, using again (7) and (8):

$$E_z y_t = \left\{ \alpha[1-(\theta_1+\theta_2)](\theta_1+\theta_2) + \frac{\theta_2}{\pi_3} \right\} \left(\pi_1 u_t(z) + \pi_2 v_t + \pi_3 \varepsilon_t\right) + \delta y_{t-1} \tag{A9}$$

Given (A9) and (A4), the conditional expectation of p_{t+1} is then found to be:

$$E_z p_{t+1} = \pi_0 + \left\{ \frac{\pi_4}{\pi_2}\theta_1 + (\pi_5+\pi_6)\left[\alpha(1-\theta_1-\theta_2)(\theta_1+\theta_2) + \frac{\theta_2}{\pi_3}\right] \right\} \left(\pi_1 u_t(z) + \pi_2 v_t + \pi_3 \varepsilon_t\right)$$
$$+ (\pi_5+\pi_6) y_{t-1} \tag{A10}$$

Substituting the expressions for $E_z p_t$ and $E_z p_{t+1}$, (A3) and (A10), into (5) and collecting terms, we get the expression for the local price level:

$$p_t(z) = \frac{(\alpha+\beta)\pi_0 + m_{t-1}}{k} + \left\{ \alpha(\theta_1+\theta_2) + \beta\left[\frac{\pi_4}{\pi_2}\theta_1 + (\pi_5+\pi_6)\left\{\alpha(1-\theta_1-\theta_2)(\theta_1+\theta_2) + \frac{\theta_2}{\pi_3}\right\}\right]\right\}$$
$$\cdot \frac{1}{k}\left(\pi_1 u_t(z) + \pi_2 v_t + \pi_3 \varepsilon_t\right) + \frac{1}{k}(u_t(z) + v_t - \varepsilon_t) + \frac{\alpha\pi_4 + \mu}{k} v_{t-1} - \frac{\delta}{k} y_{t-1}(z)$$
$$+ \frac{1}{k}(\pi_5+\pi_6)(\alpha+\beta\delta) y_{t-1} \tag{A11}$$

Equating coefficients in (A11) and (6) yields the following solutions for the undetermined coefficients:

$$\pi_0 = m_{t-1} \tag{A12a}$$

$$\pi_1 = \pi_2 = -\pi_3 \tag{A12b}$$

$$\pi_3 = -\frac{\dfrac{\beta\theta_1}{1+\beta}\mu + \dfrac{\beta\delta\theta_2}{1+\beta(1-\delta)} + 1}{k - \alpha(\theta_1+\theta_2)\left[1 - \dfrac{\beta\delta}{1+\beta(1-\delta)}(1-\theta_1-\theta_2)\right]} \tag{A12c}$$

$$\pi_4 = \frac{\mu}{1+\beta} \tag{A12d}$$

$$\pi_5 = -\frac{\delta}{k} \tag{A12e}$$

$$\pi_6 = -\frac{\delta(\alpha+\beta\delta)}{k\left[1+\beta(1-\delta)\right]} \tag{A12f}$$

Substituting the expressions for $\pi_0,\ldots,\pi_6$ into (6) and denoting $\pi_1=\pi_2=-\pi_3$ as $\pi(\mu)$ we can write the final reduced form solution for the local equilibrium price level:

$$p_t(z) = m_{t-1} + \pi(\mu)\cdot\left(u_t(z)+v_t-\varepsilon_t\right) + \frac{\mu}{1+\beta}v_{t-1} - \frac{\delta}{k}y_{t-1}(z) - \frac{\delta(\alpha+\beta\delta)}{k[1+\beta(1-\delta)]}y_{t-1} \tag{A13}$$

Aggregating (A13) over all markets yields the aggregate price level:

$$p_t = m_{t-1} + \pi(\mu)\cdot\left(v_t-\varepsilon_t\right) + \frac{\mu}{1+\beta}v_{t-1} - \frac{\delta}{1+\beta(1-\delta)}y_{t-1} \tag{9}$$

and taking expectations of (9) conditional on the information available in local markets, using (7) and (8), we obtain:

$$\mathrm{E}_z p_t = m_{t-1} + \pi(\mu)\cdot(\theta_1+\theta_2)\left(u_t(z)+v_t-\varepsilon_t\right) + \frac{\mu}{1+\beta}v_{t-1} - \frac{\delta}{1+\beta(1-\delta)}y_{t-1} \tag{A14}$$

Now, subtracting (A14) from (A13), we derive the local price surprise:

$$p_t(z) - \mathrm{E}_z p_t = \pi(\mu)\cdot(1-\theta_1-\theta_2)\left(u_t(z)+v_t-\varepsilon_t\right) - \frac{\delta}{k}(y_{t-1}(z)-y_{t-1}) \tag{A15}$$

Substituting (A15) into the local supply function (1) we derive the equilibrium level of local output:

$$y_t(z) = \alpha\pi(\mu)\cdot(1-\theta_1-\theta_2)\left(u_t(z)+v_t-\varepsilon_t\right) + \varepsilon_t + \frac{\alpha\delta}{k}y_{t-1} + \frac{(1+\beta)\delta}{k}y_{t-1}(z) \tag{A16}$$

Finally, aggregation of (A16) over all markets yields aggregate (average) output:

$$y_t = \alpha\pi(\mu)\cdot(1-\theta_1-\theta_2)\left(v_t-\varepsilon_t\right) + \delta y_{t-1} + \varepsilon_t \tag{11}$$

Appendix 2.

1. Monetary policy feedback rules.

We approached the problem of the specification of a money supply rule for each period starting from an unrestricted dynamic equation for the (log of) nominal M1, *m*, as a function of five lags of itself, of (the log of) real GNP, *y*, and of (the log of) the price level *p*. All data are quarterly, seasonally adjusted. Successive steps of reduction and reparameterization have been performed on the initial equations in order to reach a more parsimonious (and economic meaningful) representation of the monetary policy rules. The results are reported in Table A1.

There is evidence of a negative reaction of money growth to past output growth only for the postwar period, captured by the two terms in $\Delta_i y$ in equation *(2)*. The effect of past output growth on money growth in the interwar period, if any, seems to be positive. In the 1952-1968 period there is also some evidence of a negative reaction to the inflation rate. Recursive estimation of the equations over their respective sample periods and parameter constancy tests do not detect any sign of instability.

Table A1
Money growth feedback rules.

Dependent variable: Δm_t

Sample period:	*(1)* *1922(1)-1940(4)*	*(2)* *1952(1)-1968(4)*
Constant	0.0034	-0.012
	(0.0022)	(0.004)
Δm_{t-1}	0.760	0.700
	(0.103)	(0.084)
Δm_{t-4}	-0.083	0.282
	(0.107)	(0.095)
Δy_{t-1}	-0.052	-0.109
	(0.063)	(0.051)
Δy_{t-2}	0.126	-
	(0.065)	
$\Delta_2 y_{t-3}$	-	-0.148
		(0.031)
Δp_{t-1}	-	-0.272
		(0.123)
Δp_{t-3}	-0.168	-
	(0.155)	
Δp_{t-4}	0.250	-
	(0.139)	
Time	-	0.0001
		(0.00003)
R^2	0.58	0.71
σ	0.0166	0.0035
DW	1.97	1.88
AR(6) F	1.26 [0.29]	1.51 [0.19]
ARCH(6) F	0.26 [0.95]	0.17 [0.98]
CHOW(8) F	1.38 [0.23]	1.18 [0.33]

Notes: *1)* Δy, Δm and Δp are first differences of the logarithms of GDP, M1 and the GDP deflator, respectively, taken from Balke and Gordon (1986), *Historical Data*, in R. Gordon (ed.), *The American Business Cycle. Continuity and Change*, NBER, *The University of Chicago Press*, Appendix B, p.791-810. *2)* Standard errors in parentheses. *AR(6)* is a Lagrange Multiplier test for serial correlation up to the sixth order, *ARCH(6)* is the Engle (1982) test for autoregressive conditional heteroscedasticity, *CHOW(8)* is the Chow test for parameter constancy over a period of 8 quarterly observations, obtained when the equations are estimated over the 1922-1938 and 1952-1966 periods. [.] denote p-values.

2. Time-series properties of the data.

a) The postwar period: 1952-1968.

We begin by testing each variable for a unit root in (log)levels against the alternative hypothesis of stationarity around a linear deterministic trend, using augmented Dickey-Fuller (*ADF*) tests (Stock and Watson (1989), Krol and Ohanian (1990)). Results are reported in Table A2, panel A. In all cases the presence of a unit root is detected. The same tests are then repeated on the (log)differences of the variables, to ascertain the existence of a second unit root, allowing for the alternative that the series is stationary in first differences around a linear time trend: all variables are stationary in first differences.

To investigate the order of the deterministic trend with more powerful tests, the first difference of each variable is regressed against a constant, a time trend and two of its own lags. The *t*-statistic on the trend coefficient is reported, showing that only the growth rate of M1 presents a statistically significant linear deterministic trend. Omitting the time trend from the previous regression and computing the *t*-statistic on the constant provides a test for drift in the differenced variables. The last column of the table shows that only ΔR does not contain a significant drift.

In panel B of the table, the possibility that our series have common stochastic trends, displaying cointegration, is investigated. If this is the case, the number of unit roots in multivariate representations of the series is reduced and a correct first-difference specification should contain also the appropriate error-correction (stationary) terms. The omission of these terms causes misspecification of conventional *VAR* systems in first-differences. To test for cointegration we applied Johansen (1988) *trace* test (λ_{TRACE}) to systems including two (y and m), three (y, m and p), and four variables (y, m, p and R). In all cases the null hypothesis of no cointegration is not rejected, suggesting that all the multivariate specifications contain as many unit roots as variables and first-difference *VAR* systems are appropriate for estimation and inference.

b) The interwar period: 1922-1940.

The same battery of univariate and multivariate tests is applied to the interwar period. Results are reported in Table A3. As in the postwar period, all series appear to contain one unit root, whereas no evidence of a second unit root is detected. Unlike the 1952-1968 period, the quarterly growth rate of money supply does not display a linear time trend. Multivariate tests again are not able to reject the null hypothesis of no cointegration among the variables, finding no evidence of common stochastic trends.

Table A2

Tests for integration, cointegration, and time trends: 1952-1968.

A. Univariate tests.

Variable (z)	*Unit-root tests:*		*t-stat. for a regression of Δz on:*	
	ADF(z)	*ADF(Δz)*	*Time*	*Constant*
y	-0.89(4)	-5.56(3)**	1.58	4.57**
m	1.09(5)	-4.25(2)**	2.87**	3.10**
p	-0.88(3)	-3.92(1)**	1.70	2.33*
R	-3.00(4)	-4.94(4)**	0.35	0.72

B. Multivariate tests.

System	*Johansen (1988) λ_{TRACE} statistic:*
y, m	13.1
y, m, p	27.3
y, m, p, R	43.9

Notes: *1) ADF* is the augmented Dickey-Fuller test for unit roots on the following regression:

$$\Delta z_t \quad a - bz_{t-1} + c(t - \frac{T}{2}) + \sum_{i-1}^{n} d_i \Delta z_{t-i} + u_t$$

where n (reported in parentheses) is chosen to obtain white noise residuals. The null hypothesis is non-stationarity (H_0: $b=0$). Critical values are -2.93 (5%) and -3.58 (1%) for Δy and Δp, -3.50 (5%) and -4.15 (5%) for Δm, and -1.95 (5%) and -2.62 (1%) for ΔR. *2)* In Panel B Johansen's (1988) *trace statistic* for the null hypothesis of no cointegrating vectors in the system against the alternative of at least one such vector is reported. 5% critical values are: 15.4, 29.7 and 47.2 for the three systems respectively.

Table A3
Tests for integration, cointegration, and time trends: 1922-1940.

A. Univariate tests.

Variable (z)	*Unit-root tests:*		*t-stat. for a regression of* Δz *on:*	
	ADF(z)	*ADF*(Δz)	*Time*	*Constant*
y	-2.10(3)	-3.23(3)**	0.20	0.80
m	-1.20(3)	-2.45(3)*	0.72	1.23
p	-2.20(3)	-3.05(3)**	0.25	-0.37
R	-3.15(1)	-6.11(1)**	0.13	-1.11

B. Multivariate tests.

System	*Johansen (1988)* λ_{TRACE} *statistic:*
y, m	13.5
y, m, p	27.5
y, m, p, R	45.3

Note: The *ADF* test is described in the notes to Table A2. Critical values are -1.95 (5%) and -2.62 (1%) for all variables. Critical values for the λ_{TRACE} test are reported in the notes to Table A2.

References for Chapter 4

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Chapter 5

Monetary policy, credit shocks and the channels of monetary transmission. The case of Italy: 1982-1994.

1. Introduction.

Theoretical research on the microeconomics of credit markets has rapidly grown over the last fifteen years. The abandonment of the hypothesis of perfect information of all agents allowed a deeper understanding of the nature and working of financial intermediaries. In particular, the asymmetric information borrowers and (potential) lenders have on the characteristics, in terms of expected returns and riskiness, of investment projects is the basis for explaining the existence of intermediaries and, among these, of banks. The special role of banks in providing credit to agents (firms and households) who would not be able to obtain funds on the open market at acceptable terms has therefore become the focus of the literature in this field (Gertler (1988), Bernanke (1993), Kashyap and Stein (1994)).

The progress in the theory of intermediation and banking has also influenced macroeconomic analysis and especially the debate on the channels of effectiveness of monetary policy actions (Cecchetti (1994), Hubbard (1994) and Bernanke and Gertler (1995) provide selective surveys of the relevant literature). In fact, if banks perform an essential role in providing funds to agents with no alternative sources of finance, then changes in the amount of bank credit may have important effects on investment and production. This idea has been embedded in the standard *IS-LM* framework by Bernanke and Blinder (1988), who provided a simple extension of the basic macroeconomic model by explicitly assuming *imperfect* substitutability between (bank) loans and securities as both firms' liabilities and banks' assets. These modifications of the original framework (where customer-market and auction-market credit are perfect substitutes for all agents) yield an additional potential channel of monetary policy effectiveness beside the traditional *money* channel. Such conventional channel operates through changes in the quantity of banks' deposits following a change in reserves implemented by the central bank, with a resulting effect on market rates and therefore on the interest rate-sensitive components of spending. Conversely, the *credit view* of the monetary transmission mechanism focuses on the asset side of the banking sector balance sheet, stressing the possibility for the effectiveness of monetary policy actions to be enhanced if restrictions of bank credit may not be offset (at least for a significant fraction of borrowers) by the recourse to alternative sources of finance.

The empirical analysis of the relative importance of the *money* and the *credit* views of the monetary transmission mechanism has mainly concentrated on the U.S.[1], where a distinct credit channel has been found to be important by, among others, Bernanke and Blinder (1992) and Kashyap, Stein and Wilcox (1993). Only limited evidence for countries other than the United States is available (see, for example, Dale and Haldane (1993b) for an assessment of the credit channel in the United Kingdom). However, even casual observation suggests that the crucial element for the relevance of the credit channel, i.e. a low degree of substitutability between bank loans and privately issued debt, is a characteristic feature of the Italian economy. Here the financial market for privately issued debt is still very little developed, and firms rely heavily on bank loans. Following this lead, Buttiglione and Ferri (1994) have recently provided evidence of the existence of an independent credit channel of monetary policy transmission in Italy for the 1992 recession. Using micro-data for the 1968-1991 period, Rondi, Sack, Schiantarelli and Sembenelli (1993) have analyzed the response of firms of different size to monetary tightening measures. Their results show that small firms are more severely hit by monetary restrictions, confirming the relevance of agency and contract enforcement problems in the Italian context. Although not directly aimed at testing the relevance of the credit channel, the analysis of Rondi *et al.* (1993) highlights a specific mechanism whereby such a channel may operate in Italy, namely the severe reduction of external sources of funds for small firms.

The motivation of the present chapter is to provide additional results for Italy in the 1982-1994 period, where monetary policy was repeatedly used to affect the real economy. In addition we address explicitly the identification problem that is typical to analyses of this kind, formulating a set of assumptions on both the long-run and the contemporaneous relations among the variables which should allow to separate the various demand and supply disturbances hitting the economy. Our main aim is to separately consider the effects on the economy of policy shocks (deliberate monetary policy actions) and "credit shocks" originated in the bank loan market.

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[1] Bernanke (1986) and King (1986) were the first contributors to this literature (Bernanke (1993), Kashyap and Stein (1994) and Cecchetti (1994) provide assessments of the empirical findings for the U.S.).

In the remainder of this introduction we offer some brief descriptive evidence on the relevant aspects of the Italian financial structure and on monetary policy in Italy in the 1980s and 1990s. We then describe in section 2 a stylized macroeconomic framework of the Bernanke-Blinder variety, adapted to the main features of the Italian economy. Section 3 discusses the strategy of empirical investigation, implemented in section 4 by means of a small-scale *VAR* model. Section 5 summarizes the main conclusions.

1.1. The Italian case.

To illustrate the potential relevance of the credit channel of monetary transmission for the Italian economy, we shall first briefly describe the evolution of households' and firms' balance sheets in the 1980s and 1990s. Then, monetary policy and some features of the banking sector's balance sheet will be considered.

The composition of the liability side of households' and firms' balance sheets (observed every four years over the period 1980-1992) is reported in Table 1. The dependence of firms on bank credit is relatively high and rather stable throughout the whole period (accounting for 20-25% of total liabilities and for 40-55% of firms' external finance), commercial paper is virtually absent and the amount of privately issued bonds is negligible. Notwithstanding the importance of banks in financing firms, the Italian financial system is not characterized by "relationship banking", with very close bank-firm ties, which could shield firms from monetary policy actions operating through the credit channel.[2] Data on households' liability composition suggest that the importance of the credit channel as a link between monetary policy and real activity should be limited to the investment function. In fact, consumption should not be substantially affected by credit availability since Italian households do not rely heavily on credit (as shown by the extremely low ratio of total liabilities to total assets in the last row of Table 1).

The 1980s also witnessed substantial changes in financial regulations and in the conduct of monetary policy. On the institutional side, the distortionary impact of quantity controls, such as the ceiling on bank loans and the constraint on banks' portfolio, has been progressively removed. Italian monetary and financial markets expanded dramatically, gaining both in thickness and efficiency. In 1984 the interbank overnight market became fully operational; from 1988 Treasury Bills are priced on the primary market by competitive tenders, with no base price fixed by the Central Bank. By 1990 all constraints on international capital movements have been removed. Finally, in September 1990, monthly

[2] Such close bank-firm ties are distinctive features of other economies such as Japan (Hoshi, Kashyap and Scharfstein (1990)) and Germany (Cable (1985)).

rather than daily accounting of the bank compulsory reserves has been introduced and the daily volatility of the overnight rate has been drastically reduced (Angeloni and Prati (1993)). This whole reform process substantially increased the efficiency of the money market, contributing to a more effective transmission of monetary policy impulses to the financial sector.[3]

On the policy side, we observed a shift of the emphasis from quantity control to interest rate control, mainly determined by the increased importance of exchange rate targets (Angeloni and Cividini (1990)). Credit controls were abandoned in 1983, although they were reimposed for short periods, in circumstances to be considered exceptional, in 1986 and 1987. It has been argued (Buttiglione and Ferri (1994)) that direct credit controls may prevent the credit channel of monetary policy transmission to be effective. In fact, banks could use the vast amount of securities in their portfolios to shield loan supply from monetary authorities' restrictive policy. Under such circumstances asset management opens up the possibility of absorbing a policy restriction acting on the holdings of securities, thereby leaving loan supply unaltered. According to this view the removal of quantity controls and the ensuing adjustment in bank portfolios should have enhanced the importance of the credit channel of the monetary transmission mechanism from the mid-'80s onwards.

[3] Other important deregulation measures were enacted during the 1980s, which increased competitiveness and efficiency of the banking system: freedom of establishment of new credit institutions; removal of impediments to free branching; abandonment of territorial limits within which banks could operate; change in banking supervision policy from a system of case-by-case authorizations towards the application of clear and objective rules concerning capital adequacy and asset ratios.

Table 1
Liability composition of Italian Firms and Households.

Liability composition of Firms' balance sheet (percentages on total)				
Liabilities	1980	1984	1988	1992
Short term bills	0.3	0.1	0.1	0.1
Short term debt	24.6	27.4	21.4	31.2
of which:				
- *bank finance*	23.4	22.8	17.1	21.1
Medium and long term debt	19.0	19.9	15.0	15.1
of which:				
- *bank finance*	2.9	2.5	1.8	1.1
Medium and long term bonds	3.5	3.3	2.5	3.0
Shares	52.2	40.8	54.8	44.4
Other Liabilities	0.3	8.5	6.4	6.1
Total Liabilities	*100*	*100*	*100*	*100*
of which:				
- *bank finance*	26.3	25.3	18.9	22.2
Liabilities/Assets	150.1	165.9	299.9	322.4

Liability composition of Households' balance sheet (percentages on total)				
Liabilities	1980	1984	1988	1992
Short term debt	34.3	28.4	25.1	22.3
of which:				
- *bank finance*	34.3	28.4	25.1	21.7
Medium and long term debt	65.7	68.8	72.3	72.0
of which:				
- *bank finance*	25.0	27.6	29.2	29.6
Other Liabilities	0	2.8	2.5	5.7
Total Liabilities	*100*	*100*	*100*	*100*
of which:				
- *bank finance*	59.3	56.0	54.4	51.3
Liabilities/Assets	7.5	6.0	6.7	6.5

Note: Data are taken from Bank of Italy, *Annual Report*, various years.

2. *Theoretical framework.*

According to the traditional account of the monetary transmission mechanism, a contractionary policy impulse is transmitted to the banking sector through a reduction in available reserves, determining a decrease in the amount of banks' deposits. At this point, the *money* view of the transmission mechanism emphasizes the disequilibrium of agents' portfolios and the ensuing movements of bond interest rates necessary to restore equilibrium in the money market. With perfect substitutability among all financial (non-monetary) assets, called generically "bonds", investors are basically indifferent to the composition of their non-monetary portfolio, reacting only to changes in the relative quantity of "bonds" and money; moreover, firms are indifferent to the composition of their liabilities. Factors affecting only the composition of financial instruments available to the economy have no effect on aggregate demand and monetary policy effectiveness crucially depends on the absence of other liquid assets, outside the control of the central bank, acting as substitutes for banks' deposits. On the contrary, the *credit view* stresses the adjustment of banks' asset portfolio in the face of a decrease in deposits, entailing a parallel reduction of both securities and loans, given their imperfect substitutability as bank assets. With agents not able to raise funds directly on the market, the contraction of bank loans has a direct effect on spending.

This simple mechanism has been introduced in otherwise standard macroeconomic models of the *IS-LM* variety by Bernanke and Blinder (1988) and Dale and Haldane (1993a). Though with slightly different formalizations, these models reaffirm the crucial importance of two conditions for monetary policy effectiveness through a credit channel: *i)* intermediated (bank) loans and bonds issued on the market must not be perfect substitutes as sources of finance for (at least some) firms and/or households; *ii)* monetary authorities must be able to influence the supply of intermediated loans by means of changes of the level of banks' reserves or of the interest rate charged on borrowed reserves.[4] [5] The fulfilment of these two

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[4] The continuation of the process of financial innovation and deregulation which has already characterized several financial systems in the 1980s could make it more difficult the fulfilment of these two conditions in the future. The development of non-bank financial intermediaries may provide traditionally bank-dependent investors with alternative sources of finance; furthermore, allowing banks to issue liabilities with reduced reserve obligations (e.g. eurocurrency deposits and certificated of deposits) may weaken the link between reserves and loans. An assessment of these trends in the U.S. financial system and their consequences on the conduct of monetary policy is offered by Thornton (1994).

[5] In their account of the main elements of the *lending view*, Kashyap and Stein (1993) add also a third necessary condition: that some form of imperfect or sluggish price adjustment determines the non-neutrality of monetary policy actions. This condition is not specific to the validity of the channel of monetary policy effectiveness on which we focus, being necessary for any model where monetary

conditions, however, is not sufficient to ensure the existence of a credit channel of monetary transmission capable of enhancing monetary policy effectiveness. As emphasized by Hall and Thomson (1992), Dale and Haldane (1993a) and Thornton (1994), what is needed is a reaction of the rate on bank loans, following a contractionary monetary policy impulse, larger than that of the bond rate. A widening of the loan-bond interest rate spread in response to policy actions is therefore viewed as favouring the existence of an autonomous credit channel of transmission. Unfortunately, in the context of stylized aggregate macromodels, several conditions on asset demand and supply functions have to be imposed in order to ensure the presence of an operational credit channel. To illustrate this point we employ a variant of the Bernanke-Blinder and Dale-Haldane, that we take as the general theoretical framework underlying our empirical analysis, modified in accordance with the main features of the Italian economy mentioned in the previous section.

The economy is composed of four sectors (the non-bank private sector (*NBPS*), the commercial banking sector, the government sector and the central bank) and five markets (goods, credit, government bonds, banks' deposits and borrowed reserves). Our first modelling choice is suggested by the stylized facts reported in the previous section on the composition of firms' and households' liabilities: all bonds are issued by the government and the private sector obtains finance only through intermediated, non-marketable loans. This assumption meets (in an admittedly extreme way) the first requirement for the existence of a powerful credit channel for monetary policy mentioned above. Information asymmetries between potential borrowers and open-market lenders and the advantage of banks in monitoring borrowers' performance may account for the absence of debt finance. The second condition is met by assuming that loans and government bonds are not perfect substitutes as banks' assets. The difference in marketability between loans to the private sector and government bonds induces banks to hold securities, though yielding a lower return, as a buffer against unforeseen depositor withdrawals.

In more detail, the *non-bank private sector* has bank loans as the only liability and bank deposits (bearing no interest, but held for their transaction services) and government bonds as assets. For simplicity there is no cash in the model and a zero net worth is assumed. The *NBPS* balance sheet is therefore:

$$D^s + B_p^d = L^d \tag{1}$$

policy affects real variables. Therefore, the model presented, in the spirit of Bernanke and Blinder (1988) and Dale and Haldane (1993a), is cast in terms of a fix-price aggregate demand framework. In the following empirical analysis, however, movements in the price level in response to monetary policy impulses will be considered.

Loan demand is positively related to the level of aggregate demand y and a negative function of the interest rate charged on bank loans ρ:

$$L^d = L^d(y,\rho) \qquad L_y^d > 0 \quad , \quad L_\rho^d < 0 \tag{2}$$

The supply of deposits is a positive function of aggregate demand and a negative function of the interest rate on the alternative asset (government bonds) i:

$$D^s = D^s(y,i) \qquad D_y^s > 0 \quad , \quad D_i^s < 0 \tag{3}$$

The *banking* sector invests in loans to the *NBPS* and in government bonds the resources available from deposits and reserves borrowed from the central bank. In addition, banks have to reach a target balance of reserves proportional to deposits. Their (aggregate) balance sheet is:

$$L^s + B_b^d + mD^s = D^d + R^d \tag{4}$$

Banks have to decide both the total amount of disposable assets (and liabilities) and the optimal allocation of these assets between loans and securities. The demand for deposits and the (proportional) demand for reserves are positive functions of the interest rates on bonds and loans (since higher rates induce banks to increase the size of available resources), and negative functions of the rate charged by the central bank on borrowed reserves r:

$$D^d = D^d(i,\rho,r) \qquad D_i^d > 0 \quad , \quad D_\rho^d > 0 \quad , \quad D_r^d < 0 \tag{5}$$

$$R^d = mD^d(i,\rho,r) \tag{6}$$

Banks' choice of asset composition is affected by the level of bond and loans interest rates (acting through the usual income and substitution effects, the latter assumed to offset the former in the case of i). The portfolio choice also depends (through the budget constraint) on the cost of borrowed reserves. The resulting loan supply is then:

$$L^s = L^s(i,\rho,r) \qquad L_i^s < 0 \quad , \quad L_\rho^s > 0 \quad , \quad L_r^s < 0 \tag{7}$$

For both the non-bank and the banking sector, the demand for government bonds may be derived as residual from the respective budget constraints.

The balance sheet of the *central bank* is:

$$R^s = mD^s \tag{8}$$

We assume that the bank uses the interest rate on borrowed reserves as monetary policy instrument and supplies any amount of reserves commercial banks demand at the chosen rate r in order to meet their reserve target: the elasticity of reserve supply to r is therefore infinite. We have:

$$R^s = R^s(r) \qquad R^s_r = \infty \tag{9}$$

Finally, in the *goods market*, the level of real activity depends negatively on the interest rates on loans and on government bonds. In an economy where firms do not issue bonds and households have positive net assets the first relation is activated through the investment function and the second through the consumption function. The relation between interest rate on bonds and the level of activity is the result of a substitution and an income effects: if households are net lenders a rise in the interest rate increases the opportunity cost of consumption but has also a positive impact on disposable income, via a higher return on assets. By considering a negative relation between i and aggregate demand, we implicitly assume that the substitution effect offsets the income effect. We therefore have:

$$y = y(i,\rho) \qquad y_i < 0 \quad , \quad y_\rho < 0 \tag{10}$$

The equations of the model are summarized in Table 2.

Given the central bank's choice of the policy instrument r, the model may be solved for the nine endogenous variables (L^d, L^s, D^d, D^s, R^d, R^s, y, i,ρ) by imposing equilibrium in all markets together with the condition that the banks' balance sheet constraint is satisfied. Comparative statics results may then be derived by total differentiation of the equilibrium condition for the credit, deposit and goods markets, obtaining the following system:

$$\begin{bmatrix} L^d_y & (L^d_\rho - L^s_\rho) & -L^s_i \\ -D^s_y & D^d_\rho & (D^d_i - D^s_i) \\ 1 & -y_\rho & -y_i \end{bmatrix} \begin{bmatrix} dy \\ d\rho \\ di \end{bmatrix} = \begin{bmatrix} L^s_r \\ -D^d_r \\ 0 \end{bmatrix} dr \tag{11}$$

Comparative statics results for a change in the policy rate may then be easily derived. The effect of a change in r on aggregate demand is:

$$\frac{dy}{dr} = \frac{-D^d_r\left[L^s_i y_\rho + y_i\,(L^d_\rho - L^s_\rho)\right] - L^s_r\left[D^d_\rho y_i - y_\rho (D^d_i - D^s_i)\right]}{(L^d_\rho - L^s_\rho)\,(D^d_i - D^s_i) + D^d_\rho L^s_i - L^d_y\left[y_i\, D^d_\rho - y_\rho (D^d_i - D^s_i)\right] - D^s_y\left[L^s_i y_\rho + (L^d_\rho - L^s_\rho)y_i\right]} \tag{12}$$

Table 2
Theoretical framework.

Sectoral balance sheets

NBPS $\quad D^s + B_p^d = L^d$

Banks $\quad L^s + B_b^d + mD^s = D^d + R^d$

Central Bank $\quad R^s = mD^s$

Credit market

Loan demand $\quad L^d = L^d(y,\rho) \qquad L_y^d > 0 \;,\; L_\rho^d < 0$

Loan supply $\quad L^s = L^s(i,\rho,r) \qquad L_i^s < 0 \;,\; L_\rho^s > 0 \;,\; L_r^s < 0$

Deposit market

Deposit demand $\quad D^d = D^d(i,\rho,r) \qquad D_i^d > 0 \;,\; D_\rho^d > 0 \;,\; D_r^d < 0$

Deposit supply $\quad D^s = D^s(y,i) \qquad D_y^s > 0 \;,\; D_i^s < 0$

Borrowed reserves market

Reserve demand $\quad R^d = mD^d(i,\rho,r)$

Reserve supply $\quad R^s = R^s(r) \qquad R_r^s = \infty$

Goods market

$$y = y(i,\rho) \qquad y_i < 0 \;,\; y_\rho < 0$$

If $D^d_\rho\, y_i$ is sufficiently small (in absolute value) tight monetary policy decreases the level of aggregate demand. Since in our framework the effect of the government bonds interest rate i on aggregate demand works exclusively through the consumption function, it is not implausible to assume that such effect is small in magnitude.[6] The effect on the interest rate on bonds may also be derived:

$$\frac{di}{dr}=\frac{-D^d_r\left[(L^d_\rho-L^s_\rho)+y_\rho L^d_y\right]+L^s_r\left[D^s_y y_\rho-D^d_\rho\right]}{(L^d_\rho-L^s_\rho)\,(D^d_i-D^s_i)+D^d_\rho L^s_i-L^d_y\left[y_i\, D^d_\rho-y_\rho(D^d_i-D^s_i)\right]-D^s_y\left[L^s_i y_\rho+(L^d_\rho-L^s_\rho)y_i\right]} \tag{13}$$

If deposit supply and demand display sufficiently small elasticities to aggregate demand and to the loan rate respectively (i.e. if D^s_y and D^d_ρ are not too large), then i moves in the same direction as r. Also the interest rate charged by banks on loans (ρ) reacts to policy actions according to:

$$\frac{d\rho}{dr}=\frac{D^d_r\left[y_i\, L^d_y-L^s_i\right]+L^s_r\left[(D^d_i-D^s_i)-y_i D^s_y\right]}{(L^d_\rho-L^s_\rho)\,(D^d_i-D^s_i)+D^d_\rho L^s_i-L^d_y\left[y_i\, D^d_\rho-y_\rho(D^d_i-D^s_i)\right]-D^s_y\left[L^s_i y_\rho+(L^d_\rho-L^s_\rho)y_i\right]} \tag{14}$$

In order to have ρ moving in the same direction as the policy rate r it is necessary that the term ($L^d_y\, D^d_r\, y_i$) is not too large. This condition can be met by assuming either a small income effect of interest rate on consumption or a small effect of aggregate demand on the demand for banks' loans.[7]

Clearly, the effect of a monetary policy action on interest rates and demand depends on all behavioural parameters in the model. Likewise, the consequences of policy restrictions on the loan-bond spread (ρ-i) in this model depend on all demand and supply elasticities and there are no simple conditions to be imposed on them in order to sign the overall effect. However, in the original Bernanke-Blinder setup it has been shown that an additional credit channel of effectiveness of monetary policy operates only if the spread between interest rates on alternative financial instruments widens in the face of a monetary restriction (Hall and Thomson (1992), Dale and Haldane (1993a)).

Given its crucial importance, before proceeding we provide some descriptive evidence on the behaviour of the spread between the average interest rate on bank loans and the interest rate of government securities with residual life longer than one year for Italy

---

[6] If the condition referred to above holds, the determinant of the matrix of derivatives in (11) is negative and the numerator of (12) positive.

[7] Sufficiently small elasticities of deposit supply and loan demand to the level of goods demand are necessary also in the original Bernanke and Blinder (1988) model in order for interest rates to move in the same direction following a monetary policy impulse.

(Figure 1), trying to relate its movements to the monetary policy stance.

Four main peaks are visible, occurring from mid-1979 to the beginning of 1981, at the beginning of 1985, in the first half of 1986, and at the end of 1992. They correspond to four easily identifiable monetary restrictions.

In autumn 1979 monetary policy adopted a restrictive stance in response to pressures on the exchange rate: the discount rate was raised from 10.5 to 15 per cent in two steps and a ceiling was imposed on bank credit. Despite a further increase (by 1.5 percentage points) of the discount rate in 1980, inflation kept rising, also fuelled by the fast growth of the public sector deficit. On this account in 1981 monetary policy was further tightened with the imposition of a more restrictive credit ceiling on all loans in lire. In early 1982 a gradual relaxation of the restrictive stance of policy took place in reaction to the very low level of output growth. It is interesting to note that the monetary tightening is fully reflected in the spread at the end of 1979 and in 1980, but not in 1981 when the spread decreased dramatically while contractionary monetary policy measures were still in place. This phenomenon could be understood by analyzing banks' balance sheets in the 1979-1981 period. In fact this policy tightening was not reflected in a credit restriction because banks shielded loan supply from monetary policy impulses with offsetting movements of their securities holdings. This evidence is extensively commented upon by Buttiglione and Ferri (1994), who interpret it in the light of the previous expansion of banks' securities holdings caused by quantity controls (namely the portfolio constraint). According to this interpretation, the implementation of credit controls prevented the lending mechanism of monetary policy from being active. The second peak in the spread, at the beginning of 1985, follows an increase in the discount rate from 15.5 to 16.5 per cent decided by the Bank of Italy in September 1984 to curb excessive credit expansion. Such manoeuvre, reversed in January 1985, was decided during a period of steady decline in the discount rate, which, starting from a level of 19 per cent in April 1981, reached 11.5 per cent at the beginning of 1987. On the occasion of the monetary restriction of 1984, no major adjustment in banks' portfolios occurred to offset the policy measures. The third peak in the spread follows the monetary contraction enacted from January to March-April 1986 to fight devaluation expectations and speculative attacks against the lira in the foreign exchange markets. The last peak coincides with the foreign exchange market turbulence in the second half of 1992 and the exit of Italy from the *EMS*.

Though this descriptive evidence is suggestive of a link between policy actions and the loan-bond spread, it remains the possibility that the spread is influenced also by disturbances of a different nature, for example to the credit market, unrelated to monetary policy. Eventually, an answer to this question or to deeper issues such that the empirical relevance of the credit channel of monetary transmission can be obtained only after the estimation of a complete model. It seems then more fruitful to empirically estimate the model and base inference on the simulation of the estimated relationships rather then discussing theoretical implications based on a long list of assumptions on elasticities. A number of steps are necessary to deal properly with the problems involved in identification, estimation and simulation of the model. Since the outcome of the investigation is dependent on the strategy adopted to solve such problems we devote the next section to a detailed description of the methodology implemented on Italian data.

Figure 1
Spread between the banks' loan rate and the government bond rate

3. *A strategy for empirical investigation.*

Different lines of empirical research have been pursued in the literature to assess the relative importance of the various channels of monetary policy transmission and in particular the existence of an operational credit channel. Such studies, mainly applied to U.S. data, may be divided into three broad groups.

A first set of papers (Romer and Romer (1990), Bernanke and Blinder (1992), Kashyap, Stein and Wilcox (1993), among others) investigated whether the dynamic response of financial aggregates (deposits, loans and securities held by the banking sectors) and real variables (production, unemployment) to a monetary policy impulse favours a "credit channel" interpretation of the transmission mechanism or is consistent with the traditional "money channel". Bernanke and Blinder (1992), in the context of a small-scale *VAR* system, find that, in response to a positive innovation in the Federal funds rate, interpreted as a negative monetary policy shock, banks' deposits and securities contract, leaving for some months the quantity of loans unchanged. Subsequently, loans start to fall as banks' securities portfolios are being rebuilt, when also real variables react to the monetary restriction. A similar pattern of aggregate responses is interpreted as in line with the credit channel of monetary transmission. However, as noted by Romer and Romer (1990), the same dynamic responses may be consistent with the traditional money channel, if the loan reduction is viewed as the consequence of the real effects of monetary policy on production, determining a fall in credit demand. This identification problem between movements to the demand and supply of credit is explicitly addressed by Kashyap, Stein and Wilcox (1993) by comparing the behaviour of bank loans and commercial paper in firms' balance sheets following a monetary restriction. A reduction in bank loans relative to commercial paper is interpreted as evidence of a shift in the loan supply (as predicted by the credit view) and not merely reflecting the adjustment of loan demand to a contraction of production.

Focusing on interest rate dynamics, a second line of research has investigated directly the determinants of interest rates differentials, trying to establish whether the apparent predictive power of the commercial paper-government bills interest rate spread for industrial production is due to this differential being a proxy for private firms' default risk instead of signalling the monetary policy stance. Bernanke (1990) and Bernanke and Blinder (1992) favour the latter interpretation, whereas Friedman and Kuttner (1993) attribute the predictive power of the paper-bill spread to the cyclical behaviour of firms' cash flow, not directly related to monetary policy actions.

Finally, the cross-sectional implications of the credit view of the monetary transmission mechanism have been tested, among others, by Gertler and Gilchrist (1994), Oliner and Rudebusch (1995) using disaggregated data on firms' balance sheets. The fact that small firms display a sharper reduction in sales and inventory investments relative to large firms during episodes of monetary contractions is interpreted by Gertler and Gilchrist (1994) in favour of the credit channel of monetary transmission.[8] On the contrary, Oliner and Rudebusch (1995) criticize the interpretation by Kashyap, Stein and Wilcox (1993) of changes of firms' composition of external finance towards commercial paper in the face of monetary restrictions as due to the operation of a credit channel and attribute this evidence to a shift of all types of credit from small (more bank-dependent) firms to large firms, with no support for the credit view.

Our strategy of empirical investigation for Italy is close to the first set of studies mentioned above, using aggregate data, but jointly considers the dynamic movements of financial quantities and interest rates. The methodology employed is basically a Structural *VAR* (*SVAR*) (Pagan (1994) provides a recent assessment of this modelling technique and complete references). However, we pay explicit attention to a number of issues which are not usually heavily emphasized by *SVAR* modellers.

We start by imposing a probabilistic structure on the data, given by a general *VAR* model. For a generic vector of variables z, we have:

$$\begin{aligned} z_t &= A(L)z_{t-1} + v_t \\ A(L) &= A_1 + A_2 L + \dots + A_p L^{p-1} \\ v_t &\sim N(0,\Omega) \end{aligned} \tag{15}$$

Due to the nature of the time series involved, the *VAR* is likely to be non-stationary. As a consequence, the unconditional distribution of the statistical model is not defined, inference based on standard distributions cannot be applied and the autoregressive representation in (15) cannot be inverted to obtain the moving average representation. To properly deal with this issue we adopt the system cointegration analysis proposed by Johansen (1988, 1992, 1994) and reparameterize the *VAR* in (15) as follows:

[8] A similar analysis is conducted by Kashyap and Stein (1995) on banks' balance sheets, studying the differential response of loan supply to a monetary policy restriction for small and large banks. The results are interpreted as moderately in favour of the credit view of the monetary transmission mechanism.

$$
\left(I - A^*(L)L\right) \Delta z_t = A(1) z_{t-1} + v_t
$$

$$
A(1) = \sum_{i=1}^{p} A_i - I, \quad A^*(L) = A_1^* + A_2^* L \ldots + A_p^* L^{p-1}, \quad A^*_i = -\sum_{k=i+1}^{p} A_k \tag{16}
$$

As already noted in the empirical analysis of chapter 3, matrix $A(1)$ contains all long-run information about the system and its rank r (if $r<n$) yields the number of cointegration relationships among the variables. Only if the rank $r=0$ the widespread procedure of specifying the *VAR* in first differences (for a recent and often cited example see Bernanke and Blinder (1992)) would correctly allow inversion and application of standard inference without loss of relevant information. Instead, for $0<r<n$, we have $A(1)=\alpha\beta'$, where α is a n by r matrix of loadings and β is a n by r matrix of coefficients of the r cointegrating vectors. A careful treatment of cointegration is necessary in order to obtain a correctly specified *VAR* representation of the system.

The existence of a multiplicity of cointegrating vectors determines an identification problem for the parameters defining the long-run relations among the variables in the system (see chapter 3, section 3.2). A solution to it can be achieved by imposing a number of constraints on the matrix β sufficient to define it as the only matrix in the cointegrating space satisfying the constraints. A more formal condition for identification can be derived, following Johansen (1992), defining as R_i the r_i by n matrix which imposes r_i linear constraints on the ith contegrating vector:

$$
\begin{gathered}
R_i \beta_i = [0] \\
R_i \ (r_i \ x \ n), \ \beta_i (n \ x \ 1), \quad [0] \ (r_i \ x \ 1), \\
rank \ R_i = r_i
\end{gathered} \tag{17}
$$

A necessary and sufficient condition for identification can then be expressed as follows:

$$
rank \ [\ R_i \beta \ ] = r-1, \qquad i = 1, \ldots, r \tag{18}
$$

The cointegrating space is then identified when applying the restriction of one cointegrating vector to the other cointegrating vectors a matrix is obtained whose rank is equal to the total number of cointegrating vectors minus one. When the cointegrating space is identified we distinguish between the cases of just-identification and over-identification. In the latter case a χ^2 test of the validity of the over-identifying restrictions may be implemented (Johansen (1994)). A static model, like the one we sketched in the previous section, can be thought of as the long-run solution of the dynamic model fully describing the data. In principle one should then be able to identify a number of cointegrating relationships equal to the number of structural relationships posed by the model. Though difficult in practice, the analysis of the cointegrating structure of the variables is nevertheless important both as a way of checking how the data are close to the long-run structure proposed by the theoretical model and to make explicit the long-run solution of the estimated dynamic model.

Once the long-run identification problem has been addressed, a stationary representation of non-stationary series is obtained which fully describes the probabilistic structure imposed on the data, specifying a distribution for the vector of variables conditional upon the available information set:

$$\begin{gathered}\Delta z_t = \alpha\beta' z_{t-1} + A^*(L)\,\Delta z_{t-1} + v_t \\ v_t \mid I_{t-1} \sim N.I.D.(\mathbf{0}, \Sigma) \\ \Delta z_{t-1} \mid I_{t-1} \sim N.\,I.\,D.\left(\alpha\,\beta' z_{t-1} + A^*(L)\,z_{t-1}\,,\,\Sigma\right)\end{gathered} \tag{19}$$

Given the above representation of the data, we would like to derive empirical evidence on the monetary transmission mechanism by simulating the response of the system to disturbances in bank reserves, deposit supply and loan supply. To this aim, the reduced form residuals in (19) are not useful since they cannot be interpreted as disturbances to some structural relation and because, being correlated, they do not allow analysis of the response of the system to a particular shock independently from other disturbances. In order to solve both these problems we think of (19) as a reduced form representation of the following structural model:

$$\begin{gathered}\boldsymbol{A}\Delta z_t = \boldsymbol{a}\beta' z_{t-1} + \boldsymbol{B}^*(L)\,\Delta z_{t-1} + \boldsymbol{B}\boldsymbol{u}_t \\ \boldsymbol{u}_t \mid I_{t-1} \sim N.I.D.\,(\mathbf{0},\,\boldsymbol{I})\end{gathered} \tag{20}$$

The structural residuals are thought of as orthogonal to each other and the specification of the parameters in the matrices $\boldsymbol{A}$ and $\boldsymbol{B}$ allows some structural interpretation. The specification of the cointegrating relationships, $\beta' z_{t-1}$, is the same in the structural and in the reduced form of the system. The following restrictions ensure that (19) is derived from (20):

$$\boldsymbol{A}v_t = \boldsymbol{B}\boldsymbol{u}_t, \quad \alpha = \boldsymbol{A}^{-1}\boldsymbol{a}, \quad \boldsymbol{A}^*(L) = \boldsymbol{A}^{-1}\boldsymbol{B}^*(L) \tag{21}$$

A short-run identification problem arises if (20) is to be the unique structural model from which (19) is derived or, equivalently, if $\boldsymbol{u}$ is the unique set of unobservable structural shocks that can be associated with the observed reduced form residuals v. The solution to this problem lies with the imposition of a sufficient number of constraints on the parameters in $\boldsymbol{A}$ and $\boldsymbol{B}$. The reduced form provides us with $n(n+1)/2$ estimated elements in the variance-covariance matrix Σ so that at most $n(n+1)/2$ parameters can be estimated in the matrices $\boldsymbol{A}$ and $\boldsymbol{B}$. Formal analysis (Giannini (1992), Hamilton (1994)) provides necessary, and necessary and sufficient, conditions for identification. Also in this case we can have exact-identification and over-identification and derive a test for the over-identifying restrictions. Although this approach has been proposed some time ago (Bernanke (1986), Blanchard and Watson (1986)), many *VAR* models used to analyze the monetary transmission mechanism are just-identified, imposing a diagonal $\boldsymbol{B}$ and a lower-triangular $\boldsymbol{A}$ with ones on the principal diagonal. This identification method, originally introduced by Sims (1980) does not allow to fully exploit theory to set restrictions and to test for over-identification restrictions. In this context, the only contribution theory can give is on the ordering of variables in the *VAR*, which obviously affects the derivation and interpretation of the shocks (recently, Gordon and Leeper (1994) have stressed the importance of using theory to impose restrictions on the $\boldsymbol{A}$ and $\boldsymbol{B}$ matrices).

Imposing identification restrictions on the simultaneous feedbacks among the variables included in the *VAR* implies that the higher is the frequency of observation the easier should be the solution to the identification problem. From this perspective, the use of monthly data is clearly advisable. Moreover, the inspection of the correlation matrix of the reduced form residual may be used in association with theory in order to derive identifying restrictions: the observation of zero correlations between some reduced form residual might be informative on the plausibility of different structures for the $\boldsymbol{A}$ and $\boldsymbol{B}$ matrices. Finally, as in the case of the long-run parameters, testing for the validity of the over-identifying restrictions may give some support to the chosen structure as a valid explanation of the data.

Having constrained the simultaneous feedbacks and the long-run response of the system one can compare its behaviour with the prediction of the theory by looking at the full dynamic adjustment process, i.e. by analyzing the full dynamic response of the system to the relevant structural disturbances. Inversion of the structural error-correction model (20) is necessary to achieve this result. In order to do so, we rewrite the structural model as follows:

$$
\begin{aligned}
G(L)z_t &= A^{-1}B\,u_t \\
G(L) &= \Big((I-L)\left(I - A^*(L)\,L \right) - \alpha\beta' L \Big)
\end{aligned}
\tag{22}
$$

By multiplying both sides of (22) by the adjoint of $G(L)$, $G^a(L)$, we obtain:

$$
G^a(L)\cdot G(L)\,z_t = G^a(L)\cdot A^{-1}B\,u_t \tag{23}
$$

where $G^a(L) \cdot G(L)$ is the determinant of $G(L)$. If we define $d(L)=[detG(L)]/(1-L)$ we can exclude the presence of unit root from the determinant of $G(L)$ and rewrite (23) as follows:

$$d(L)\ (1-L)\ z_t = G^a(L)\ A^{-1}B\ u_t \tag{24}$$

d(L) does not contain unit roots and can be inverted to derive the impulse response functions, yielding the reaction of any variable in the system at time $t+s$ to any structural disturbance hitting the system at time t. In fact we have:

$$\Delta z_t = \frac{G^a(L)}{d(L)} A^{-1}B\ u_t = G_0 u_t + G_1 u_{t-1} + \ldots + G_s\ u_{t-s} + \ldots \tag{25}$$

The analysis of the dynamic response of the system completes our strategy of empirical investigation.

In the light of the rather general framework developed above we now move to the discussion of the empirical application proposed in this chapter with our baseline model given by the simple theoretical framework outlined in section 2. We impose a particular structure on this model and estimate a dynamic version of it, including an equation for inflation to model the supply side of the economy. The variables included in the system are:

RP : interest rate on Bank of Italy's repurchase agreement operations;
LYD : (log of the) seasonally adjusted index of industrial production;
INFL : annual inflation of the consumer price index;
LOAN : (log of) real bank loans;
DEP : (log of) real bank deposits;
RL : average interest rate on bank loans;
RB : average interest rate on government bonds with residual life longer than one year.

The final form of the structural model we will estimate is the following:

$$
\begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & a_{43} & 1 & 0 & 0 & 0 \\ 0 & 0 & a_{53} & 0 & 1 & 0 & 0 \\ a_{61} & 0 & a_{63} & a_{64} & 0 & 1 & a_{67} \\ a_{71} & 0 & a_{73} & 0 & a_{75} & 0 & 1 \end{bmatrix}
\begin{bmatrix} \Delta RP_t \\ \Delta LYD_t \\ \Delta INFL_t \\ \Delta LOAN_t \\ \Delta DEP_t \\ \Delta RL_t \\ \Delta RB_t \end{bmatrix}
- \alpha\beta' \begin{bmatrix} RP_{t-1} \\ LYD_{t-1} \\ INFL_{t-1} \\ LOAN_{t-1} \\ DEP_{t-1} \\ RL_{t-1} \\ RB_{t-1} \end{bmatrix}
+ B^*(L) \begin{bmatrix} \Delta RP_{t-1} \\ \Delta LYD_{t-1} \\ \Delta INFL_{t-1} \\ \Delta LOAN_{t-1} \\ \Delta DEP_{t-1} \\ \Delta RL_{t-1} \\ \Delta RB_{t-1} \end{bmatrix}
$$

$$
+ \begin{bmatrix} b_{11} & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & b_{22} & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & b_{33} & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & b_{44} & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & b_{55} & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & b_{66} & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & b_{77} \end{bmatrix}
\begin{bmatrix} u^r \\ u^y \\ u^{infl} \\ u^{ld} \\ u^{dd} \\ u^{ls} \\ u^{ds} \end{bmatrix} \qquad (26)
$$

Our choice of the policy variable (the repo rate) is in accord with other studies on Italian interest rates (see, among others, Ansuini, Fornasari and Paruolo (1992)), but differs from the choice made in other recent papers (Buttiglione and Ferri (1994)), where the overnight rate has been given the role of indicator of the monetary policy stance[9]. With reference to interest rates it is also worth noting that the assets underlying *RL* and *RB* have comparable duration, so that their relative movements do not capture term structure changes. As far as our indicator of real activity is concerned, it may be argued that the share of national product explained by industrial production has declined over time. However, recent studies on the Italian economy support the validity of industrial production as a cyclical indicator (Schlitzer (1993)). Moreover, such variable may appropriately capture the effect of credit restrictions on investment. This indicator of real activity is the only variable we have included in the *VAR* in seasonally adjusted form. We believe that this choice is justified by the peculiar pattern of the seasonally unadjusted index, which for Italy shows a very strong "August effect", when a large fraction of Italian firms interrupt production. The data are monthly over the sample 1982(6)-1994(12). Earlier observations have been excluded because the developments of the monetary and financial markets in the 1980s limit the comparability with data from earlier periods, and because data on *RP* are not available before 1982.

Six of the seven equations in (26) are dynamic versions of the equations listed in

[9] We note that these two rates move very closely to one another and that preliminary investigation suggests that our results are robust to the substitution of the repo rate with the overnight rate.

Table 2: the equations describing the credit market, the deposit market and the equations for reserves supply and for the goods market. We have excluded an equation for reserve demand assuming that the only shock relevant to this equation is the deposit demand shock, which can be identified by the other equations in the system. An equation for the inflation rate is included because, although some price stickiness is a necessary condition for any monetary policy to work, it does not seem sensible to exclude a supply side from any estimated model. The inclusion of the inflation rate in the system allows also to address the "price puzzle" observed by Sims (1992) in a multi-country study of the effects of monetary policy (see also Eichenbaum (1992)). The puzzle consists in a perverse response of the price level to a monetary contraction: following a positive innovation in short-term interest rates, signalling monetary *tightening*, the price level appears to *increase* in France, Germany, Japan and the U.K.; only for the U.S. the response of the price level is negative, though after a considerable lag. A prolonged period of inflation following a monetary contraction is a finding which is difficult to rationalize by any existing business cycle theory. Sims' own explanation relies on the possibility that monetary authorities decide policy tightening on the basis of information on future inflation not captured by past behaviour of the variables analyzed. If this is the case, the observed increase in interest rates reflects the effort of the monetary authorities to combat future inflationary pressures and the subsequent increase in inflation is not a perverse response to monetary tightening, but measures the portion of price pressures not avoided by the enacted contractionary policy. Eichenbaum (1992) interprets this finding as casting serious doubts on the validity of using interest rate innovations as indicators of monetary policy disturbances.[10] Given the open debate in the literature, it seems important to address the issue directly for Italy including a measure of price movements in the system.

Identification of the relevant shocks is obtained by having a sufficient number of variables in the model and by imposing some short-run restrictions on the contemporaneous feedbacks. As the empirical literature mentioned at the beginning of this section has clearly illustrated, trying to establish evidence on the relevance of the credit view using for instance only a three-variable *VAR* for deposits, loans and a measure of output is not a meaningful exercise since it does allow identification of asset demand from asset supply shocks. A mix of prices and quantities in the estimated model seems to be a necessary condition to achieve identification of demand and supply shocks (Friedman and Kuttner (1993)). Within our

[10] Another possible explanation of the puzzle could be a direct effect of the increase in borrowing costs for firms, transferred onto sales prices, or the presence in the price index used of mortgage payments and other items directly linked to interest rate levels.

seven-equation model identification is then achieved by imposing restrictions on the simultaneous feedbacks between shocks, i.e. on the parameters of the matrices $\boldsymbol{A}$ and $\boldsymbol{B}$. The latter matrix is assumed diagonal and the b_{ii} parameters deliver the estimated standard error of each equation and allow standardization of residuals.

Some explanation of the identifying restrictions we propose on $\boldsymbol{A}$ is in order. The shocks to loan and deposit demand, u^{ld} and u^{dd} respectively, are identified by ruling out any contemporaneous effect from all variables but inflation on loan and deposit demand. This choice is inspired and empirically sustained by the results obtained in structural modelling of Italian money demand (see the results reported in chapter 3). Loan supply is assumed to react simultaneously to all interest rates and inflation; deposit supply shocks are identified by allowing contemporaneous feedback between the government bond rate, the policy rate and the quantity of deposits. Any contemporaneous feedback is ruled out for the policy rate, the index of activity and inflation. This is a stringent set of assumptions which over-identifies the short-run parameters with twelve over-identifying restrictions, whose validity will be tested in the empirical section.

Equation (26) does not specify any long-run restriction. Under the null of correct specification of the model it seems reasonable to expect six cointegrating relations to be delivered in the case of a non-stationary system: loan demand, loan supply, deposit demand, deposit supply, aggregate demand and aggregate supply. Given that the problem of the identification of the number of cointegrating vectors is totally separable from the problem of the identification of the parameters in the cointegrating vectors we condition our choice of long-run identification restrictions on the results of the cointegration analysis and describe them in the empirical section. Having proposed a solution to the long-run and short-run identification problems we will then implement impulse response analysis to describe the dynamic adjustment of the system to the identified structural disturbances.

4. *The econometric evidence.*

We begin our empirical investigation by setting up a seven-equation *VAR* system. The dimension of the *VAR* satisfies a necessary condition to obtain identification of asset demand and supply shocks. We then evaluate whether the estimated reduced form system provides a satisfactory representation of the data generating process through a battery of diagnostic tests. Next, we take up cointegration analysis and identification of the long-run relationships and then move to structural modelling by imposing a set of short-run identifying restrictions and testing their validity. Finally, some evidence on the dynamic adjustment of the model is provided by means of impulse response functions and forecast error variance decompositions.

4.1 The estimation of the system.

The estimated system is specified with a lag of order five; the deterministic part includes a constant, a linear trend and seasonal dummies. Although the industrial production variable is already in seasonally adjusted form, both deposits and (to a lesser extent) loans show seasonal patterns requiring the introduction of a full set of dummies. Diagnostic tests on the reduced form residuals show that the presence of few outliers has a sizeable effect on the residual normality in the equations for inflation and output. Such observations are concentrated in August of various years (1982, 1983, 1984 and 1987), suggesting a seasonal pattern not adequately captured by the seasonal adjustment used. Four point dummies are therefore included in the system to obtain a normal distribution of the reduced form residuals. The same procedure is followed for the 1992(1) observation, which is responsible for the non-normality originally detected in the deposit equation. The system is then estimated with this set of point dummies included and the relevant results are summarized in Table 3. Panel B of the table reports single-equation tests for autocorrelation, normality and autoregressive conditional heteroscedasticity of the residuals.[11] There is no evidence for heteroscedasticity and the autocorrelation tests, though significant at the 5% level in three cases, do not reveal major problems. The break-point recursive stability Chow test on the system (Figure 2) reveals some marginal sign of instability only at the end of 1992, due to the sharp movements in all interest rates during the *EMS* crises. Normality of residuals seems to be a problem in the equations for the policy rate *RP* and the bank loan rate *RL*. The presence of several outliers in the interest rate equations is a difficult problem to handle. In fact non-normality could raise serious problems with the application of Johansen's maximum likelihood procedure, requiring residual normality for efficient estimation. However, it must be noted that, as pointed out by Gonzalo (1994), when the Johansen's estimator is compared with alternative estimators for the cointegrating vectors, it displays more desirable properties (at least in large samples), even when the *VAR* residuals are drawn

[11] The tests are implemented at the equation level rather than at the whole system level because the size of the system does not allow a sufficient number of degrees of freedom for the system tests. For a general discussion of both single-equation and system diagnostic tests see Doornik and Hendry (1994a, 1994b).

Table 3
Reduced form estimation.

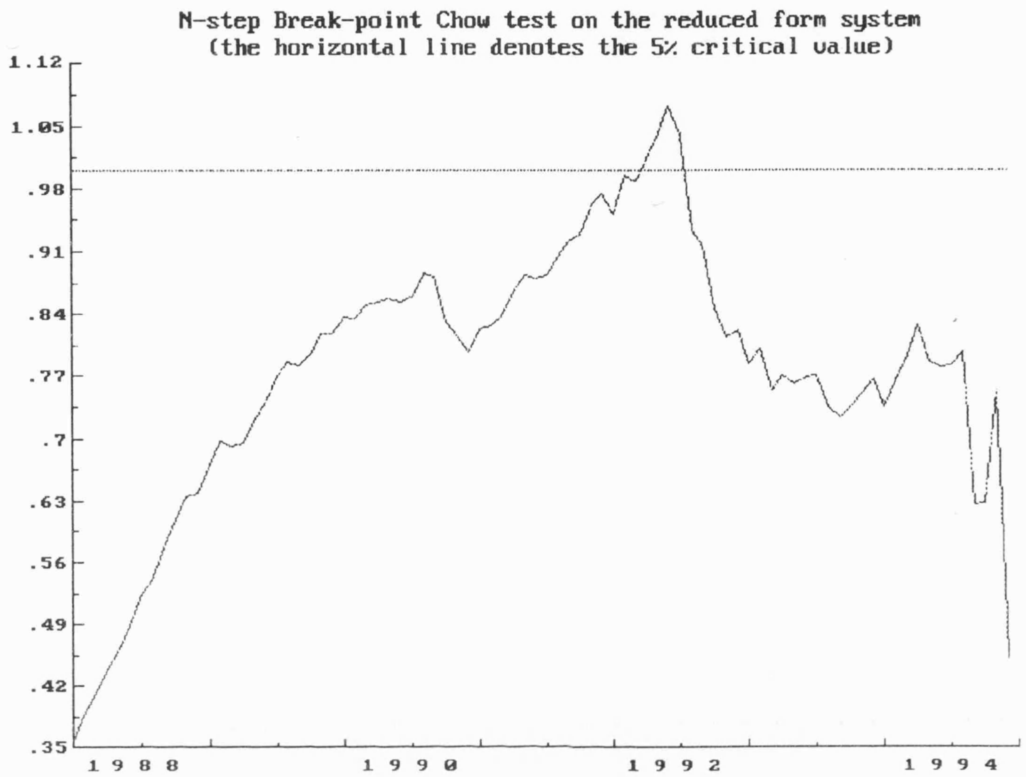
A) *Correlations of Unrestricted Reduced Form residuals*

	RP	*RB*	*RL*	*LOAN*	*DEP*	*LYD*	*INFL*
RP	1						
RB	0.38	1					
RL	0.47	0.56	1				
LOAN	0.19	0.03	-0.10	1			
DEP	-0.04	-0.09	-0.21	0.13	1		
LYD	0.01	0.16	-0.06	-0.09	0.06	1	
INFL	0.03	-0.01	0.09	-0.18	-0.13	0.03	1

B) *Residual mis-specification tests on reduced form equations*
(p-values in parentheses)

Equation for:	σ	*AR 1-6* $F(6,92)$	*Normality* $\chi^2(2)$	*ARCH 7* $F(7,84)$
RP	0.913	0.25 (0.96)	12.63 (0.00)	1.88 (0.08)
RB	0.345	2.35 (0.04)	3.86 (0.14)	1.32 (0.25)
RL	0.169	3.13 (0.01)	41.11 (0.00)	1.40 (0.22)
LOAN	0.012	2.73 (0.02)	0.38 (0.83)	1.14 (0.35)
DEP	0.009	0.65 (0.69)	3.84 (0.15)	0.97 (0.46)
LYD	0.017	1.29 (0.27)	2.72 (0.26)	0.58 (0.77)
INFL	0.239	0.37 (0.90)	2.82 (0.24)	0.37 (0.92)

Figure 2
Break-point system Chow stability test.



from non-normal distributions. Moreover, removing interest rate outliers with dummies may not be appropriate for economic policy analysis, because outliers might capture very significant and decisive moves by the monetary authorities. Therefore, we decided to keep these observations, after checking that the cointegration results do not change substantially when point dummies are introduced to eliminate the largest outliers.

Our reduced form describes a closed economy. It might be rightly argued that this is acceptable for the U.S. but not for Italy. In this light, we tried to augment the model with the inclusion of the real effective exchange rate but did not find any evidence for the significance of such variable. In particular, Granger-causality tests could reject the hypothesis of some additional predictive power of exchange rate movements for all variables in the system.

Panel A of Table 3 reports the correlation matrix of the unrestricted reduced form residuals. This is informative since it reveals that high simultaneous correlation seems to be limited to interest rates and inflation. The within-period relations between prices and quantities seem to be limited to some effect of inflation on loans and deposits. We intend to exploit this information in order to impose some testable over-identifying restrictions on the short-run identification scheme implemented in the final stage of the analysis.

4.2 Cointegration and long-run identification.

Results from the application of Johansen's (1988) Full Information Maximum Likelihood procedure for cointegration are summarized in Table 4. We report the results of the usual two tests for cointegration: the maximum eigenvalue and the trace statistics, with appropriate critical values as computed by Osterwald-Lenum (1992). We also consider a small sample correction, obtained by replacing, in the computation of the statistics, the number of observations T by the difference between T and the product of the length of the *VAR* lag, M, times the dimension of the *VAR*, N (in our case $T\text{-}MN=151\text{-}35=115$). Such a correction is proposed and discussed by Reimers (1992). The evidence from the cointegration analysis points clearly towards non-stationarity of the system but it is not unequivocal on the number of cointegrating vectors. Using the corrected statistics there is evidence of one or two cointegrating vectors; given the difficulties we encountered in the economic interpretation of a second cointegrating vector, we decided to proceed under the assumption of one valid cointegrating relationship in the system.

Some long-run identification restrictions are then imposed on the cointegrating vector, assuming a long-run relation among the loan rate, the inflation rate and the deviation of industrial production from a linear trend. This hypothesis imposes four over-identifying restrictions on the vector and may be tested by means of a likelihood ratio test. This is implemented in panel B of Table 4, where the estimated coefficients on the restricted vector are reported. A negative effect of the bank loan rate and a positive effect of inflation on the deviation of industrial production from trend are detected. The test of the over-identifying restrictions does not reject the assumed form of the vector. We report in Figure 3 the unrestricted and restricted cointegrating vectors.

The long-run analysis does not provide any evidence on the monetary transmission mechanism although the long run solution of our model features one of the necessary conditions for the credit channel to be operational, namely the relation between real activity and the interest rate on bank loans. To shed some further light on the monetary transmission mechanism and the importance of the credit market we revert to short-run identification and simulation.

Table 4
Cointegration analysis and long-run identification.

A) *Cointegration analysis*
(r denotes the number of valid cointegrating vectors)

Hypothesis:	$r=0$	$r\leq 1$	$r\leq 2$	$r\leq 3$	$r\leq 4$	$r\leq 5$	$r\leq 6$
λ_{MAX}	44.9	41.6	36.6	19.4	10.9	8.5	4.9
(without correction)	(58.5)	(54.2)	(47.6)	(25.3)	(14.2)	(11.1)	(6.4)
95% crit.value	49.4	44.0	37.5	31.5	25.5	19.0	12.2
λ_{TRACE}	167	122.0	80.3	43.8	24.4	13.4	4.9
(without correction)	(218)	(159)	(105)	(57.0)	(31.7)	(17.5)	(6.4)
95% crit.value	146.8	114.9	87.3	63.0	42.4	25.3	12.2

B) *Restricted cointegrating vector*

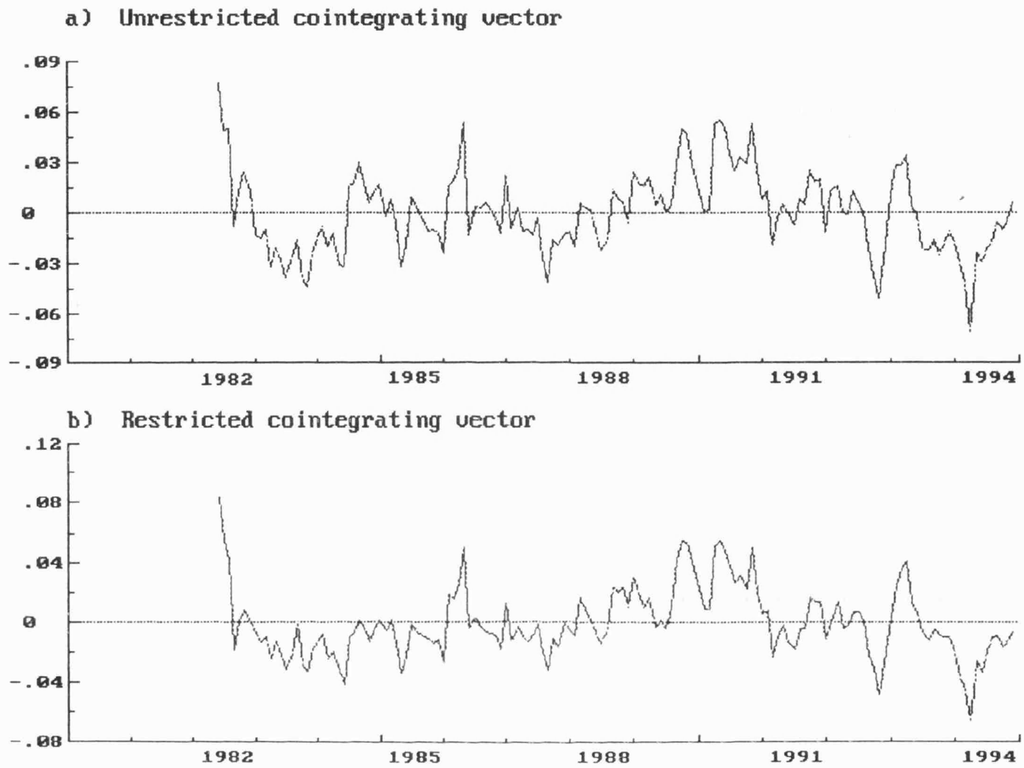
Restricted cointegrating vector:
*(first vector normalized on LYD; * denotes restricted coefficients)*

RP	*RB*	*RL*	*LOAN*	*DEP*	*LYD*	*INFL*	*Trend*
0^*	0^*	-0.0202	0^*	0^*	-1	0.010	0.00098

LR test of restrictions: $\chi^2(4) = 7.28$ *(p-value: 0.12)*

Identified long-run relation: $LYD = -0.0202\ RL + 0.010\ INFL + 0.00098\ Trend$

Figure 3
Unrestricted and restricted cointegrating vector
(normalized on LYD)



4.3 *Short-run identification and structural VAR analysis.*

The assumptions on the contemporaneous relationships among the disturbances to the *VAR* equations used to identify structural shocks have been briefly described at the end of the previous section. We start the structural *VAR* analysis by estimating the system subject to the set of twelve over-identifying restrictions on the matrix **A**, as shown in equation (26).

Loan demand shows a barely significant negative effect of the inflation rate, whereas the average interest rate on loans reacts significantly, with a positive sign, to all interest rates in the system, though the reaction to the policy rate is quantitatively small (*RL* increases by about 5 basis points in response to an increase in *RP* by about 80 basis points). No feedback from the quantity of loans to the loan rate (the element a_{64}) is detected, pointing towards an infinitely elastic within-month loan supply curve. As expected, there is a significantly negative simultaneous effect from inflation on deposit demand (captured by the element a_{53}), while the interest rate on government bonds reacts contemporaneously to the policy rate (a_{71}) but is not significantly affected by inflation (a_{73}) and by the volume of deposits (a_{75}). Reduced form innovations coincide with structural form innovations for the policy rate, output and inflation.

Given this set of results, we proceeded to a further estimation of the structural *VAR*, constraining to zero three of the least significant coefficients in the previous estimate (a_{64}, a_{73} and a_{75}) in order to obtain more efficient estimates. Table 5 shows the final estimates of the coefficients in the **A** and **B** matrices. To facilitate the interpretation of the simulation results we report here the inverted **A** matrix, capturing the simultaneous effects of the structural shocks on the variables in the system ($\epsilon_i = b_{ii}u_i$). The restrictions are not rejected by a likelihood ratio test at the 5% confidence level and the final estimate of **A** includes only significant coefficients.

$$
\begin{aligned}
v_{RP} &= \epsilon_{RP} \\
v_{Y} &= \epsilon_{Y} \\
v_{INFL} &= \epsilon_{INFL} \\
v_{LOAN} &= -\,\mathbf{0.004}\ \epsilon_{INFL} + \epsilon_{LOAN} \\
v_{DEP} &= -\,\mathbf{0.007}\ \epsilon_{INFL} + \epsilon_{DEP} \\
v_{RL} &= \mathbf{0.072}\ \epsilon_{RP} + \mathbf{0.061}\ \epsilon_{INFL} + \mathbf{0.181}\ \epsilon_{RB} + \epsilon_{RL} \\
v_{RB} &= \mathbf{0.139}\ \epsilon_{RP} + \epsilon_{RB}
\end{aligned}
$$

We note that the policy rate affects contemporaneously both the bond rate and the loan rate, but the within-month reaction of the former is almost twice as large. Moreover,

Table 5
Short-run identification and Structural VAR analysis.
(sample period: 1982(6)-1994(12))

Final identification restrictions on A matrix

Parameter estimates of A and B matrices			
Matrix element	Coefficient	St. error	t-value
a_{43}	0.004	0.003	1.45
a_{53}	0.007	0.003	2.69
a_{61}	-0.047	0.014	-3.42
a_{63}	-0.061	0.040	-1.53
a_{67}	-0.181	0.036	-4.95
a_{71}	-0.139	0.028	-4.89
b_{11}	0.799	0.046	17.38
b_{22}	0.025	0.001	17.38
b_{33}	0.254	0.014	17.38
b_{44}	0.009	0.001	17.38
b_{55}	0.008	0.001	17.38
b_{66}	0.125	0.007	17.38
b_{77}	0.278	0.016	17.38

LR test of the over-identifying restrictions: $\chi^2(15) = 24.8$ *(p-value = 0.06)*

The estimated VAR in structural form is the following:

$$\begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & a_{43} & 1 & 0 & 0 & 0 \\ 0 & 0 & a_{53} & 0 & 1 & 0 & 0 \\ a_{61} & 0 & a_{63} & 0 & 0 & 1 & a_{67} \\ a_{71} & 0 & 0 & 0 & 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} \Delta RP_t \\ \Delta LYD_t \\ \Delta INFL_t \\ \Delta LOAN_t \\ \Delta DEP_t \\ \Delta RL_t \\ \Delta RB_t \end{bmatrix} = \alpha\beta' \begin{bmatrix} RP_{t-1} \\ LYD_{t-1} \\ INFL_{t-1} \\ LOAN_{t-1} \\ DEP_{t-1} \\ RL_{t-1} \\ RB_{t-1} \end{bmatrix} + B^*(L) \begin{bmatrix} \Delta RP_{t-1} \\ \Delta LYD_{t-1} \\ \Delta INFL_{t-1} \\ \Delta LOAN_{t-1} \\ \Delta DEP_{t-1} \\ \Delta RL_{t-1} \\ \Delta RB_{t-1} \end{bmatrix}$$

$$+ \begin{bmatrix} b_{11} & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & b_{22} & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & b_{33} & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & b_{44} & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & b_{55} & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & b_{66} & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & b_{77} \end{bmatrix} \begin{bmatrix} u^r \\ u^y \\ u^{infl} \\ u^{ld} \\ u^{dd} \\ u^{ls} \\ u^{ds} \end{bmatrix}$$

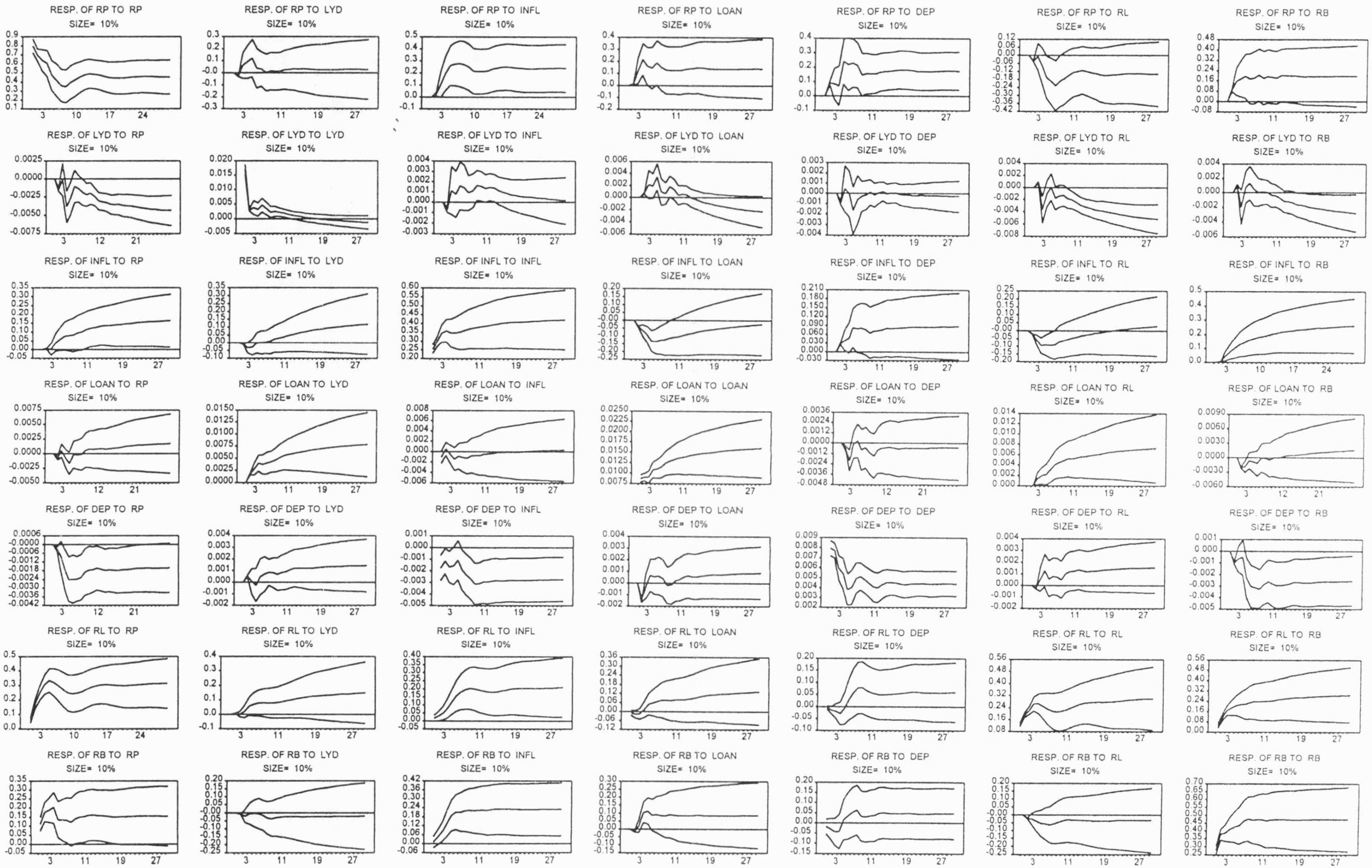
the loan rate reacts simultaneously to disturbances to the bond rate. As expected, inflation has a negative contemporaneous impact on the demand for deposits and a positive, though relatively small, effect on the loan rate.

The (not statistically rejected) imposition of the final set of short-run identifying assumptions allows the identification of shocks to loan and deposit demand and supply functions, to the policy rate, and to output and the rate of inflation. In order to analyze the dynamic response of the system to the individual disturbances we can invert the cointegrated *VAR* to obtain interpretable impulse response functions[12]. The whole set of impulse response functions is reported in Figure 4. Each column plots the response of all seven variables to a specific shock in the system. Such responses are displayed for the log-levels of the quantity variables and the levels of the interest and inflation rates. Point estimates along with 90% confidence intervals (computed by maximum likelihood following Giannini (1992) and Hamilton (1994)) are shown up to thirty months after the shocks. The impact effect is determined by the short-run identifying restrictions whereas the long-run response is shaped by the cointegrating relationships, with the previously tested restrictions on one cointegrating vector imposed on the system.

The first column reports the dynamic response of the system to a monetary policy tightening, i.e. a shock to *RP*. With a lag of few months output declines significantly, providing evidence in favour of some effectiveness of monetary policy. Inflation positively responds to the tightening yielding some evidence of a (not quantitatively important) "price puzzle" for Italy. Interest rates on both government securities and bank loans react positively: after an initial stronger reaction of *RB*, the response of *RL* becomes larger and the loan-bond spread does indeed widen following a monetary tightening episode. The second column shows the response to a disturbance in *LYD*, interpreted as an aggregate demand shock. The lack of response of the inflation rate suggests a rather flat aggregate supply curve, a necessary condition for monetary policy effectiveness; loans and deposits gradually increase over time and, among interest rates, only the bond rate displays a negative reaction to the shock. A disturbance to *INFL* (third column) has no significant effect on output and loans, while deposits show a prolonged decline, following the negative contemporaneous effect. All interest rates are positively (and significantly) affected, with quantitatively very similar responses: the loan-bond interest rate spread does not open up following an unexpected movement in inflation. The fourth column gives the responses to a shock to *LOAN*, interpreted as a loan demand disturbance. The interest rate on loans tends

[12] In doing so we rule out non-fundamental representations for the process generating the residuals. For a discussion of this point see Hansen and Sargent (1991) and Lippi and Reichlin (1993).

Figure 4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to rise, giving support to the identification assumption adopted. Also the policy rate tends to rise (at least initially), possibly as a consequence of a monetary authorities' reaction to potential inflationary pressures fuelled by credit expansion. However, inflation does not show any significant positive response. Output initially displays a positive (though not strongly significant) response, which turns negative after about one year from the loan shock. The reaction of a shock to deposit demand, *DEP*, reported in the fifth column, features a plausible positive response in inflation and in the policy rate, witnessing a monetary tightening following a money demand shock, and no reaction of output. It could be noted at this point that if a more limited system including financial quantities only were estimated, the results obtained as response to the *LOAN* and *DEP* shocks could be, probably wrongly, interpreted as evidence against the relevance of credit disturbances and the lending channel of monetary transmission.

The last column shows the reaction to a shock in the interest rate on government bonds, *RB*. The response of output is negative, though not highly significant, and inflation does rise somewhat (perhaps showing some ability of market rates in anticipating future inflation). The negative response of deposits is in line with the interpretation of the disturbance to *RB* as a negative shock to deposit supply and limits the separation between prices and quantities to the simultaneous feedback. Finally, the sixth column reports responses to a shock to *RL*, interpreted as a loan supply disturbance. The response of output is negative and quantitatively large, supporting the importance of credit supply shocks. There is no evidence of a "price puzzle" in response to an increase in the interest rate on bank loans. The interest rates on government bonds and the policy rate do not show a significant reaction, while there is a marginally significant positive response of the quantities of loans. This last result conflicts somewhat with our interpretation of shocks to *RL* as being loan supply disturbances. However this anomaly could be caused by the irregular behaviour of bank loans following the removal of the ceiling on bank loans.

To supplement the evidence provided by the impulse response analysis we report, in Figure 5, the results from the forecast error variance decomposition (*FEVD*). For example, considering the case of industrial production, the *FEVD* indicates what proportion of the error variance the econometrician makes in predicting industrial production can be attributed to the shocks identified as structural for the other variables of the system. Therefore, if interest rates are significant in explaining industrial production after the transmission of the monetary stance to the real economy, we then expect that the variance of the innovations in these variables explains an increasing share of the variance of the prediction error for industrial production as the forecasting horizon increases.

We report in each row of Figure 5 the point estimates of the *FEVD* up to a thirty-month horizon along with 90% confidence intervals. The analysis of the *FEVD* confirms the importance of interest rates in explaining industrial production. We note that the share of the forecast error variance explained by the own shocks constantly declines for industrial production starting from 100% in the one-period ahead forecast (due to our short-run identifying assumptions) to reach a share of 25% in the thirty-period ahead forecast. The policy rate and the interest rate on long bonds explain respectively 20% and 25% of the thirty-period *FEVD* of industrial production, whilst about 12% of the same variance is explained by the shocks in the interest rate on bank loans. A very small share of the *FEVD* in industrial production is explained by shocks to demand of loans and the demand in deposits, independently from the time horizon chosen. The *FEVD* for the other variables confirms the tendency of the interest rates to move together, the policy rate being a crucial element in explaining the behaviour of other rates with limited feedback effects. If we consider quantities, we note the importance of inflation and the interest rate on bonds in explaining deposits, while the share of the variance of the *FEVD* in bank loans depending on its own shocks remains high and stable as the time horizon increases.

To sum up, several tentative conclusions may be drawn from the above results:

i) monetary policy actions, captured by unexpected movements in the repo rate, have a non-negligible effect on industrial production. Autonomous disturbances to loan supply ("credit shocks") have a quantitatively important effect on industrial production, confirming the relevance of the banking sector as a source of finance for firms in the Italian economy. Although innovations in the repo rate and autonomous disturbances to loan supply are orthogonal by construction in the whole sample, in one relevant episode (in the occasion of the *EMS* crisis in 1992) sizeable positive shocks in both rates are observed over a short time span (Figure 6);

ii) no evidence of a quantitatively important perverse price response to monetary policy tightening is detected;

iii) the response of bank deposits to policy contractions is significantly negative, whereas loans do not show any dynamic reaction: this evidence is difficult to interpret as supporting the credit view of the monetary transmission mechanism;

iv) there is evidence of a widening of the loan-bond interest rate spread in response both to policy shocks and credit supply shocks. Therefore this differential contains information on both sources of disturbances and cannot be uniquely associated to the stance of monetary policy.

Figure 5

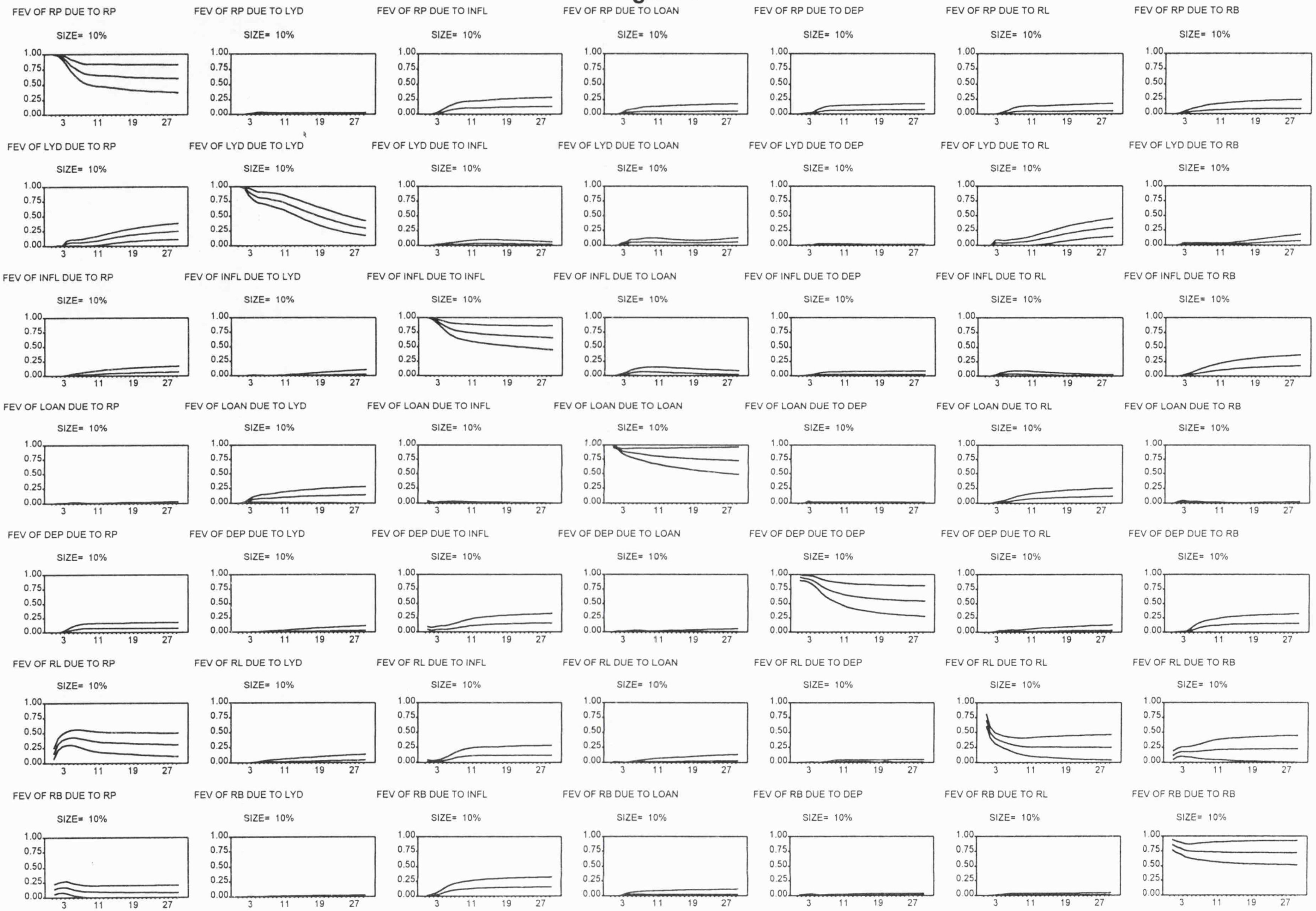

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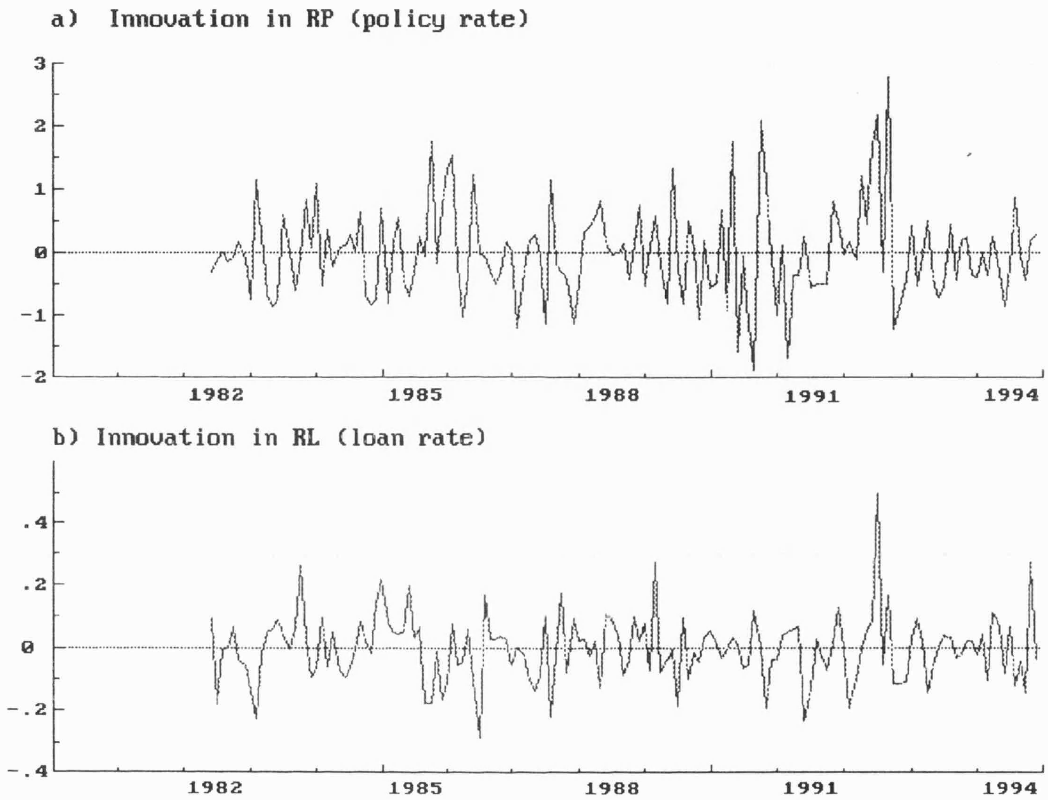
Figure 6
Innovations in the policy and loan rates



5. Conclusions.

This chapter aimed at providing some preliminary evidence on the relevance of the credit channel of transmission of monetary policy impulses for the Italian economy and the importance of disturbances to the bank loan market. A basic theoretical framework, derived by the simple Bernanke-Blinder setup, has been adapted to the Italian case and used to provide some guidelines for the design of the estimation strategy. The fundamental problem of the identification of disturbances of a different nature has been directly addressed within the structural *VAR* modelling technique, applied to a seven-variable system including three relevant interest rates (the policy rate, the bond rate and the bank loan rate), two financial quantities (bank loans and deposits), the industrial production index and the inflation rate.

Estimation and simulation of the system, with a set of (data-admissible) restrictions on both the long-run and the contemporaneous relations among the variables provides a series of results for the 1982-1994 period. Monetary policy is effective, though perhaps more through the traditional deposit channel than through an autonomous lending channel, whereas disturbances to credit supply have an even more pronounced effect on output. The loan-bond interest rate spread shows a positive reaction not only to monetary policy contractions, but also to credit supply shocks and inflation does not show any perverse response to monetary tightening.

Although the preliminary nature of our investigation suggests caution in interpreting the results, the overall picture emerging from the analysis, though not supporting a "credit view only" of the monetary transmission mechanism, suggests that bank loan supply disturbances have played a distinct and non negligible role in determining fluctuations in real variables in Italy over the 1980s and the early 1990s.

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Chapter 6

The response of consumption to income: the case of anticipated tax changes.

1. Introduction.

One of the main implications of the rational expectations-permanent income (*REPI*) model of consumption is that current consumption should incorporate all information on future income and interest rates available to individuals. Innovations in consumption should therefore reflect only unanticipated changes in real lifetime resources and not predictable income variations. Starting from the classic paper by Hall (1978) these implications of the *REPI* model have been subjected to a thorough econometric investigation. Overall, there is a substantial body of evidence seriously challenging the empirical validity of the model.

On the one hand, assuming stationarity of the labour income process, Flavin (1981) concluded that the response of aggregate consumption to actual income is too strong to be consistent with the underlying *REPI* model: consumption exhibits *excess sensitivity* to the anticipated component of income movements. On the other hand, Deaton (1987) and Campbell and Deaton (1989) found that the empirical observation that consumption is smooth relative to fluctuations in observed income -traditionally interpreted as evidence in favour of the permanent income hypothesis- is inconsistent with the *REPI* model. If labour income is characterized by a difference-stationary process with positively autocorrelated first differences, the *REPI* model implies that consumption should be more -and not less- volatile than income. The observed behaviour of consumption displays *excess smoothness* with respect to innovations in (permanent) income. As shown by Deaton (1992) and Flavin (1993), the two results of excess sensitivity and excess smoothness of consumption obtained in time-series studies are intimately related.[1]

[1] It is important to note that the inconsistency between the implications of the theory and the data referred to above concerns the Hall-Flavin version of the permanent income *cum* rational expectations hypothesis. In fact, as observed by Falk and Lee (1990), this version of the model is substantially different from the original formulation of the permanent income hypothesis in Friedman (1957) and

More recently, some efforts have been made in order to reconcile the implications of the *REPI* model with the available evidence. For example, Quah (1990) provides an explanation for excess smoothness in consumption based on agents' different reactions to permanent and transitory movements in labour income. A joint explanation for excess smoothness and excess sensitivity is offered by Pischke (1991), who assumes that agents ignore information on *aggregate* income (which is available only with a lag and is not very informative on the behaviour of *individual* income) but react optimally to their own income process. Finally, Caballero (1990b) highlighted the potential of the precautionary saving motive in providing an explanation for the observed excess sensitivity (if lagged income changes are positively correlated with the expected income variance, which, under the precautionary saving hypothesis, determines the consumption path) and excess smoothness (if a positive correlation is allowed between innovations in the level and variance of income).

Instead of conducting traditional time-series analyses, some authors have pursued a different research strategy and provided evidence against the *REPI* hypothesis by studying the response of consumption under "natural experiments", i.e. clearly identified "*income shocks with predictable and well-understood effects on future income*" (Poterba (1988, p.413)). Fiscally-induced income changes are primary candidates in this respect. Examining episodes of explicitly temporary income tax changes, Poterba (1988) found that U.S. consumption reacts to such temporary tax shocks by more than predicted by the *REPI* hypothesis; moreover, consumers do not appear to respond to tax announcements of future changes in tax policy. Wilcox (1989) studied the impact of pre-announced increases in U.S. social security benefits on aggregate consumption expenditure: his results show a strong effect on consumption, especially on durables, at the time when the increases were paid.

The analysis performed in the present chapter follows this line of research, exploiting the time lag between the announcement of changes in income taxation in the United Kingdom (usually made in the March-April and Autumn Budget Statements) and the slightly delayed enactment of such fiscal measures. If the Ricardian Equivalence proposition holds, there should be no detectable effect on spending when tax changes are implemented, since, with government expenditure held fixed, they should be perceived only as a change in the timing of taxation and not in the overall present value of tax liabilities. Moreover, even if the Ricardian proposition is not correct but the standard formulation of the *REPI* hypothesis

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also from the Muth (1960)-Sargent (1979) rational expectations version of it. In particular, both these earlier formulations of the permanent income hypothesis do not yield the same implications in terms of sensitivity and smoothness of aggregate consumption that have been challenged by much recent empirical research.

is valid, consumption expenditure should not be affected by the implementation of tax changes, since they should have already been reflected in spending at the announcement date. Therefore, a detectable response of consumption to tax changes implementation may be interpreted as valuable evidence against the *REPI* model and the Ricardian Equivalence proposition.

The chapter is organized as follows. In section 2 a brief presentation of the relevant empirical literature is provided and the interpretation of the test is discussed. Section 3 describes the data and the specification of the test; some methodological issues are also addressed. Section 4 contains the empirical results and section 5 the main conclusions.

2. *Relevant literature and interpretation of the test.*

Several episodes of changes in income tax and transfer policy in the United States over the last three decades have been considered as "natural experiments" useful to test models of consumption behaviour. The response of consumption to explicitly temporary income tax changes and to the implementation of pre-announced fiscal measures has been the focus of the empirical analysis, since the basic *REPI* model predicts a limited reaction of consumption to temporary disposable income movements and no reaction of current consumption to previously announced income changes.

The effects on consumption of the 1968 surtax (a temporary increase in personal income tax) and the 1975 tax rebate (coupled with other temporary decreases in taxes and increases in transfer payments) received special attention in the empirical literature, albeit with sometimes conflicting results. Modigliani and Steindel (1977), using traditional consumption function estimates, found that the 1975 rebate had only a modest impact on spending, in line with the implications of the permanent income and life-cycle theories. Analyzing the same episode, Blinder (1981) estimated a marginal propensity to consume out of a temporary tax cut larger than that implied by the permanent income theory for pure windfall gains, but smaller than the impact on spending of a permanent tax reduction. In contrast with Blinder's estimates, Blinder and Deaton (1985), examining both episodes (1968 and 1975), concluded that consumers did not spend on the basis of the temporary changes in their disposable income, their behaviour being broadly consistent with the REPI model. Finally, Poterba (1988), using higher-frequency (monthly) data and adopting the modern consumption Euler equation approach for testing, documented a positive response of consumption on nondurables to the 1975 rebate.

A more consistent pattern of results is obtained when clearly pre-announced changes in tax policy are analysed. Both Blinder and Deaton (1985), studying the 1981-1984 promised tax reductions, and Poterba (1988), examining several episodes from 1964 to 1986, found that consumption did not react to the announcement of future tax changes. Moreover, Poterba and Summers (1987), in a detailed case study of the 1981 tax cut -announced well in advanced and only gradually implemented-, showed that consumption expenditure was not affected by the announcement of tax cuts, whereas both overall spending and spending on nondurables responded positively to the implementation of tax cuts. The magnitude of the estimated elasticities of consumption expenditure to disposable income imply that expenditure on durables reacted most to tax cuts implementation.[2] Finally, Shapiro and Slemrod (1995) studied the response of consumers to a change in income tax withholdings, altering the timing of agents' income receipts without affecting their lifetime resources, occurred in the U.S. in 1992: some 43% of the surveyed consumers manifested the intention of spending most of the extra take-home pay, revealing a behaviour in contrast with the *REPI* hypothesis.

Although the study of specific examples of changes in fiscal policy yields valuable information on consumption behaviour, the analysis of *recurrent* episodes of this kind over a long period of time may provide a more powerful test of the *REPI* model. Wilcox (1989) estimated the effect of changes in social security benefits -implemented at least six weeks after announcement- on aggregate U.S. spending for the 1965-1985 period. His results show that total retail sales strongly react to social security benefits increases with a long-run elasticity of around 0.20, mainly due to the durable expenditure component (with an elasticity of 0.40).

Following these lines, in the present chapter we exploit some characteristic features of the British political system, allowing for a precise dating of the announcement of fiscal measures. In fact, changes in income tax and allowances are announced in the annual Budget Statement (occurring in March or April, with only few exceptions[3]) but are implemented only with a lag of between two and five months, the main reason being that the Inland Revenue needs some months to adjust all taxpayers' *PAYE* tax codes to the change in income taxation. The Budget Statement is a much publicized event, widely covered by the media and not only by the financial press. This is an ideal set-up for testing whether consumption

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[2] The estimated elasticities are between 0.10 and 0.15 for total consumption and around 0.04 for nondurables consumption (all estimates are statistically significant).

[3] In some years, additional Budgets were announced in October-November.

behaviour is consistent with the basic *REPI* hypothesis. Sumner (1991) carried out a first study of U.K. data, concentrating on a short period (1976-1988) and showing that expenditure on non-food items does react to changes in income at the implementation dates, with an elasticity around 0.4.[4]

We provide a more extensive analysis of the U.K. experience, under at least three respects. Firstly, we study a longer sample period (1960-1990): results obtained using a sample going back to 1960 may be interpreted more confidently as reflecting a behavioural regularity, whereas those derived from the 1976-1988 period only may be substantially affected by few episodes of sizeable tax cuts implementation (especially in 1979); moreover, in the earlier part of the period, Budget announcement very often resulted in unanticipated income tax *increases* whereas the post-1975 period displays a prevalence of income tax *cuts*.[5] This feature may be important in order to discriminate between alternative explanations for the failures of the *REPI* hypothesis reported by Poterba, Wilcox and Sumner, usually attributed to two potential causes: liquidity constraints or myopia. Secondly, we employ data for three sub-categories of consumption goods, characterised by a different degree of durability, which enables us to be more precise as to what kinds of consumption expenditure react most to anticipated income changes. Finally, we try to control for a number of variables (expected real interest rates, relative price movements, unanticipated news about consumers' real income and wealth), whose omission from the analysis could make the interpretation of the results more questionable.

Even though the *REPI* hypothesis correctly characterises the consumers' decision process in the absence of constraints beyond the intertemporal budget constraint, the presence of imperfections in credit markets may not allow individuals to increase their consumption expenditure, after the announcement of a reduction in income taxes but before the tax cut is actually implemented. If such liquidity constraints affect a substantial part of the population, the aggregate effect may well be the lack of response of consumption expenditure to announcements of future increases in disposable income. However, liquidity constraints would not prevent immediate downward adjustment of consumption after announcements of future tax *increases*. This asymmetric response of spending distinguishes

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[4] This estimate is obtained when only the continuing effect of implemented tax changes (not including the rebate payment due to the delay in implementation) is considered. The elasticity to the rebate component is around 0.1.

[5] The unanticipated component of Budget announcements is obtained by eliminating the change in income tax attributed to (anticipated) allowances indexation to past inflation. Details on this point are provided in the next section and in the *Appendix*.

the presence of liquidity constraints from a more radical departure from the assumptions of the *REPI* model, i.e. myopic behaviour. In this event consumers do not behave as rational, forward-looking agents, basing instead consumption decisions on the level of *current* (and not *permanent*) income. Consequently, current consumption should display no reaction to announcements of future tax changes of either sign. Allowing for a different consumption response to announcements of tax cuts and tax increases may help the interpretation of the detected implementation effect.

The finding of a response of consumption to the implementation of tax changes may also be viewed as evidence against the validity of the Ricardian Equivalence proposition. The most fundamental version of this proposition states that, with government spending held fixed, decreases in lump-sum taxes should not have any real effect, since rational agents would increase their savings in response, anticipating offsetting future tax increases. Deviations from Ricardian Equivalence may have various explanations, first of all the non-lump-sum nature of taxes, creating distortions with real effects.[6] Moreover, even ruling out distortions (for example assuming an inelastic labour supply), consumption may respond to income tax reductions if the certainty equivalence principle does not hold (because of a non-quadratic utility function) and agents accumulate precautionary savings. In this case, the change in the timing of taxation will reduce agents' income uncertainty, causing a decrease in precautionary savings (Barsky, Mankiw and Zeldes (1986)). Developing this idea, Kimball and Mankiw (1989) show that the announcement of a future tax cut causes an immediate increase in consumption, followed by further increases until the tax cut is actually implemented.

However, although either the non-lump-sum nature or the insurance effect of the income tax system may explain the real effect of tax changes, the result that consumption reacts to pre-announced income tax changes at the (delayed) *implementation* date contradicts the very basic assumptions of the Ricardian Equivalence proposition. Therefore, the results obtained by Poterba (1988), Wilcox (1989) and Sumner (1991) may be confidently interpreted as strong evidence against one of the tenets of the "neoclassical view of fiscal policy" (Barro (1989)).

[6] Non-lump-sum taxes are analysed in all the empirical studies surveyed above, with perhaps the only exception of Wilcox (1989). In fact, at least for the individuals already receiving social security benefits, their amount is predetermined by past wage history and increases in benefits come closer to the definition of (negative) lump-sum taxes than other forms of taxation.

3. *The specification of the test and the data.*

3.1. *Testing framework and data analysis.*

Our test is based on a log-linear specification of the first-order condition (Euler equation) from the standard optimization problem of a representative consumer, endowed with rational expectations (Hall (1978, 1988), Hansen and Singleton (1983), Abel (1990), Deaton (1992)). Consider an infinitely-lived consumer choosing the optimal path of consumption to solve the following problem[7]:

$$\max_{C_{t+i}} \; E_t \sum_{i=0}^{\infty} \left[\frac{1}{1+\delta}\right]^i U(C_{t+i}) \tag{1}$$

subject to the budget constraint:

$$W_{t+i} = W_{t+i-1}(1+r_{t+i}) + Y_{t+i} - C_{t+i} \qquad \textit{for all } i \geq 0 \tag{2}$$

and the transversality (no Ponzi-game) condition:

$$\lim_{i\to\infty} \; E_t W_{t+i} \prod_{k=1}^{i} \left[\frac{1}{1+r_{t+k}}\right] \tag{3}$$

where δ is the time-invariant rate of time preference, C_t is consumption in period t, W_t is wealth at the beginning of period t, r_t is the real interest rate between period t-1 and t, Y_t is labour income in period t, and E_t denotes (rational) expectations formed on the basis of the information set available in period t, I_t. The first-order necessary condition for the above problem is:

$$U'(C_t) = \frac{1}{1+\delta} E_t\left[U'(C_{t+1})(1+r_{t+1})\right] \tag{4}$$

Assuming that $U(.)$ is of the constant relative risk aversion (*CRRA*) class, e.g.:

$$U(C_t) = \frac{C_t^{1-\rho}}{1-\rho} \tag{5}$$

where ρ is the coefficient of relative risk aversion, the Euler equation (4) becomes, after rearranging terms:

[7] We assume here intertemporal separability of the utility function. Generalizations of the above framework, allowing for non-separability between consumption and leisure and for the possibility that government expenditure may be a substitute for private expenditure are provided by Mankiw, Rotemberg and Summers (1985) and Bean (1986).

$$(1+\delta) = E_t\left[\left[\frac{C_{t+1}}{C_t}\right]^{-\rho}(1+r_{t+1})\right] \equiv E_t(X_{t+1}) \tag{6}$$

Letting $c \equiv ln\ C$ and $x \equiv ln\ X$, the distributional assumption needed to obtain a log-linear form of (6) is that Δc and r are generated by a covariance stationary Gaussian process (Hansen and Singleton (1983))[8]. Under this assumption x_{t+1} is conditionally normal with mean μ_t and variance σ^2. Therefore:

$$E_t(X_{t+1}) = exp\left[E_t(x_{t+1})+\frac{\sigma^2}{2}\right] \tag{7}$$

Combining (6) and (7) and rearranging (using $c_{t+1}=E_tc_{t+1}+\epsilon_{t+1}$ and the approximations $ln(1+\delta)\simeq\delta$ and $ln(1+r)\simeq r$) we obtain:

$$\begin{aligned}\Delta c_{t+1} &= \frac{1}{\rho}\left[\frac{\sigma^2}{2}-\delta\right] + \frac{1}{\rho}E_t r_{t+1} + \varepsilon_{t+1}\\ &= \qquad \alpha \qquad + \phi\ E_t r_{t+1} + \varepsilon_{t+1}\end{aligned} \tag{8}$$

where ϵ_{t+1} is orthogonal to all variables known at t or earlier. The rate of change of consumption is positively related to the level of the expected real interest rate.[9] (8) is the log-linear form of the Euler equation on which our empirical analysis is based.

We use monthly, seasonally unadjusted, data for the volume of retail sales of three different categories of consumption goods -Food, Clothing and Footwear, Household Durable Goods- taken as representatives of the broader categories of non-durable, semi-durable, and durable goods, and data for total retail sales (all items). Original retail sales value indices were deflated using the corresponding indices of retail prices. A seasonally adjusted series for the retail sales volume index has also been used in the analysis. A complete list of the variables used and their sources is reported in the *Appendix* (section C).

Table 1, panel A, shows basic descriptive statistics for the monthly rate of change in real consumption expenditure over the whole sample period (1959(10)-1990(9) for the unadjusted data; 1960(11)-1990(9) for the All Item adjusted series). To have an idea of the

[8] An alternative rationalization for the log-linear form of the Euler equation (mentioned by Wilcox (1989)) assumes a different functional form for $U(.)$, such that marginal utility is a linear function of the percentage deviation from bliss-point consumption, and a fixed real interest rate.

[9] With an intertemporally separable utility function, the coefficient ϕ measures both the degree of intertemporal substitution and the degree of risk aversion ($\phi=1/\rho$). Hall (1988) and Attanasio and Weber (1989) discuss these interpretations of ϕ and the possibility of separating the elasticity of intertemporal substitution from (the reciprocal of) the coefficient of relative risk aversion.

importance of (deterministic) seasonal variability of consumption, the table also reports the standard error of a regression of the monthly rate of change of consumption expenditure on a complete set of monthly dummy variables and of monthly trending seasonals (as in Muellbauer (1983) and Sumner (1991)), to capture demographic trends and changes in seasonal patterns over time.[10] As the results show, the residual variability of the rate of change of expenditure on durable and semi-durable goods (around 4%) is higher than that of expenditure on nondurables.

The stochastic properties of the error term in (8) are crucial in assessing the validity of the underlying theory. In particular, in the basic *REPI* model, the error term should be orthogonal to all past information; therefore residual serial correlation should not be detected. In fact, all equation residuals display a very high degree of serial correlation, as shown by the large values of the Box-Pierce Q statistic for residual serial correlation up to the 24th order. Several theoretical justifications for serially correlated errors have been offered in the consumption literature, including time aggregation, the effect of transitory consumption, non-separabilities in the utility function and durability. Time aggregation and the existence of transitory consumption would introduce a first-order moving average component in the error term, whereas durability and (other forms of) utility function non-separability could generate possibly more complex error structures.[11] Even though we are not specifically interested in explaining the nature of this feature of the data, it seems worthwhile to investigate briefly the form of such serial correlation, since its presence can determine our choice of the estimation technique.

When an *MA(1)* error process is added to the previously estimated equation for the rate of change of consumption expenditure, the estimated MA coefficients are all negative and highly statistically significant, with point estimates ranging from -0.37 to -0.72 (Table 1, panel B). The substantial drop in the value of the Q statistic shows that the *MA(1)* term captures the bulk of serial correlation. In two cases (Household Durables and All items - adjusted data) the value of Q is below the 10% critical level, whereas for the other three series (especially for the expenditure on Food) some sign of residual serial correlation is still

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[10] In the equation for Household Durable goods two additional dummy variables have been included to take care of outliers which substantially affected the normality of residuals. They take the value of *+1* in 1965(9) and 1975(4), when expenditure increased by 47% and 40% respectively, and *-1* in 1965(10) and 1975(5), when expenditure decreased by 31% and 54% respectively. The 1975 episode may be due to an announced increase in the Value Added Tax on durables.

[11] Time aggregation would generate a positive sign of the *MA(1)* coefficient, whereas transitory consumption would yield a negative coefficient. However, if the transitory element of consumption is uncorrelated across individuals, its presence should not affect the behaviour of aggregate consumption series.

Table 1
Data analysis

A) Descriptive statistics on consumption expenditure series.

Monthly rate of change of cons. exp. on:

	Food	*Clothing and footwear*	*Household durables*	*All items*	*All items (adj. data)*
Mean	0.08	0.27	0.25	0.07	0.20
St dev.	7.39	19.27	9.66	11.94	1.71
σ	1.66	4.24	4.11	1.97	-
DW	2.90	3.00	2.68	2.86	2.79
Q(24)	190.3	153.7	91.4	136.6	87.8

Notes: The sample period is 1959(10)-1990(9) (1960(11)-1990(9) for the All Items adjusted data series). Means, standard deviations and σ are expressed in percentage points. σ, *DW* and *Q(24)* are the standard error, the Durbin Watson statistic, and the Box-Pierce statistic for residual serial correlation up to the 24th order from a regression of the monthly rate of change of real consumption expenditure on a complete set of monthly dummy variables and of monthly trending seasonals. In the equation for Household Durables two additional *+1/-1* dummy variables have been introduced in 1965(9)-1965(10) and in 1975(4)-1975(5) to take care of outliers. For the All items (adjusted data) series the regression includes only a linear time trend and a *+1/-1* dummy variable in 1975(4)-1975(5). The *Q* statistic is distributed as a χ^2 with 24 degrees of freedom on the null hypothesis of no serial correlation. Critical values are: 36.5 (5%) and 43.0 (1%).

B) Estimates of Euler equations with MA(1) errors.

Monthly rate of change of cons. exp. on:

	Food	*Clothing and footwear*	*Household durables*	*All items*	*All items (adj. data)*
MA(1)	-0.53	-0.72	-0.37	-0.51	-0.49
	(0.054)	(0.054)	(0.054)	(0.054)	(0.053)
σ	1.46	3.40	3.86	1.74	1.33
DW	2.02	2.07	2.02	2.04	1.99
Q(24)	60.4	46.6	31.5	42.9	26.5

Notes: Sample period as in panel A. This part of the table reports the estimated coefficients of an *MA(1)* error process (standard errors in parentheses), added to the equations estimated in panel A of the table.

Table 1/contd.

C) Tests for non-separabilities in the utility function.

Dependent var.: Monthly rate of change of expenditure on:

F-test of four lags of monthly rate of change of expenditure on:	*Food*	*Cl.&Foot.*	*Hous.Dur.*
Food	-	1.33	0.22
Clothing & Footwear	1.80	-	0.46
Household Durables	0.61	1.84	-
σ	1.45	3.39	3.96
Q(24)	43.6	45.3	38.5

Notes: The 5% critical value for the *F* test with (4,333) degrees of freedom is 2.40. The sample period is 1960(1)-1990(9).

detected. Therefore, before proceeding, we briefly investigate whether non-separability or durability may be responsible for this feature of the data.

The rationale for serial correlation of the Euler equation disturbance term which has recently spurred much theoretical and empirical work is the presence of durable goods, yielding a flow of services for several periods, and purchased only infrequently by consumers. The implications for expenditure on durables of the basic (frictionless) version of the *REPI* model was originally provided by Mankiw (1982). Extending Hall's (1978) original framework to durable goods (in particular assuming separability between durables and nondurables), Mankiw showed that the Euler equation disturbance should follow an *MA(1)* process, with a negative coefficient equal, in absolute value, to one minus the rate of depreciation of durables. However, the empirical analysis of U.S. postwar quarterly data showed that the disturbance term in the equation for durables expenditure had almost white noise properties, strongly rejecting the *REPI* model for durables. Subsequent research extended the basic *REPI* model for durables in several directions. Startz (1989) showed that ignoring the existence of (quadratic) costs of adjusting the stock of durables may lead to serial correlation in the error term in addition to the usual *MA(1)* component. Bar-Ilan and Blinder (1988), as an alternative to the stock-adjustment model, assumed lumpy transactions costs for durables and derived an (S,s) decision rule for durables purchases. In their model, changes in permanent income might lead to a very large response of durables expenditure, with additional effects lasting for several periods. Again, neglecting this source of dynamics might lead to the detection of a serially correlated disturbance term in the Euler equation for durables. Finally, Caballero (1990a) allowed for slowness in the response of a fraction of consumers to news, generating a high-order moving average representation of the process for the rate of change of consumption expenditure. A sharp difference then arises in the time-series behaviour of durables and nondurables. The sum of the autocorrelations of changes in expenditures should be positive and close to zero for nondurables, and negative and decreasing in the case of durables, reflecting a negative (and not very distant from -1) sum of the *MA* coefficients. This extension of the *REPI* model is potentially useful in explaining both the presence of some serial correlation, even after allowing for an *MA(1)* component, in the error term of the Clothing and Footwear and All Items Euler equations, and the relatively small *MA(1)* coefficient estimated for Household Durables expenditure.[12] However, when an *MA(12)* model for the Euler equation disturbance term is estimated for

[12] If interpreted according to the simple formulation of the *REPI* model for durables proposed by Mankiw (1982), an *MA(1)* coefficient of -0.37 would imply an implausibly high monthly depreciation rate of Household Durables of 0.63.

Household Durables and Clothing and Footwear, the sum of the estimated *MA* coefficients does not show the reversion towards -1 implied by the slow adjustment hypothesis. Indeed, this sum is -0.73 (with a standard deviation of 0.169) for Clothing and Footwear and -0.35 (0.161) for Household Durables, both extremely close to the previously estimated *MA(1)* coefficients.[13] Therefore, the slow response hypothesis does not seem capable of explaining the empirical behaviour of our durables expenditure series.[14]

The neglect of non-separabilities in the utility function both over time and across different categories of goods may be another explanation for the serially correlated pattern of residuals derived from simple Euler equations such as (8). If the utility function is not separable across goods and over time, the marginal utility of consumption of a particular good will depend on the current and past levels of consumption of that good and of other goods. In our testing framework, the resulting Euler equation for consumption of good i will contain as regressors also lagged rates of change of expenditure on good i and on other categories of goods. Therefore, in order to assess the empirical importance of non-separabilities of that kind, we augmented the basic Euler equations (8) for our three sub-categories of consumption goods in turn with four lags of the dependent variable and four lags of the rate of change of expenditure on the other two sub-categories and test for the joint significance of the latter blocks of regressors. The results, in the form of F-tests, are reported in Table 1, panel C, together with basic statistics on the augmented equations. There is no evidence of non-separabilities across different goods categories: in all cases the F-test does not reject the null hypothesis of separability at the 5% significance level. The sharp reduction in the values of the Q statistic, if compared with those of the basic Euler equation in panel A of the table, is entirely due to the presence of four lagged values of the dependent variable. The pattern of coefficients on these regressors (all negative and declining towards zero in absolute value) is consistent with the negative first-order *MA* coefficient in the Euler equation error term reported in panel B of the table.

Summarising the above discussion, for all consumption expenditure series we found strong evidence of residual serial correlation in the estimation of simple Euler equations.[15]

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[13] In the case of expenditure on Clothing and Footwear, allowing for a high order *MA* error term removes the residual serial correlation still present in the Euler equation with *MA(1)* errors.

[14] In addition, this hypothesis cannot apply to the nondurables (Food) series, given the negative and highly statistically significant *MA(1)* coefficient and the still negative value of the sum of the *MA(12)* coefficients: -0.28 (st. dev. 0.165).

[15] This result is consistent with those reported by Wilcox (1989) for the U.S. and by Sumner (1991) for the U.K. Wilcox attributes serial correlation to non-separabilities of the utility function over time and across goods and adopts the augmented Euler equations used above (table 1, panel C)

Allowing for a first-order moving average disturbance term completely eliminates serial correlation for the Household Durable goods and the All items (adjusted data) series. For the remaining series, there is some evidence of a more complex error structure, not easily attributable to a slow and gradual adjustment of consumption expenditure to permanent income news or to non-separabilities over time and across goods in the utility function. This evidence, possibly due to non-deterministic seasonality effects, will be taken into account in the adopted estimation procedure.

3.2. Measures of the announcement and implementation effects and estimation methodology.

To analyse the response of consumption expenditure to the announcement and subsequent implementation of income tax changes, we include in the basic Euler equation (8) a set of variables measuring the effect of tax changes on consumers' disposable income, based on the full year effect on tax revenue of changes in income taxation estimated by the Treasury and published in the *Financial Statement and Budget Report* (*FSBR*, 1960-1990). The formulation of the *REPI* hypothesis under test requires the construction of a variable measuring the perceived effect of tax changes on disposable income at the *announcement* dates, reflecting only the *unexpected* variation in consumers' real income prospects. Since periodic changes in nominal allowances were enacted throughout the sample period, *de facto* providing some form of allowance indexation to past inflation, some part of the announced income tax changes reflects *predictable* adjustments of disposable income to the past inflation rate. Therefore, only the residual *unpredictable* part of the announced changes should be expected to have some effect on consumption expenditure under the *REPI* hypothesis. The *Appendix* (section A) explains in detail how an estimate of the monthly percentage unexpected variation in disposable income perceived at Budget announcements was constructed. The resulting variable, denoted by *ANN*, is plotted in Figure 1, a positive value corresponding to a tax cut. With the notable exception of the 1981 Budget, unanticipated tax increases are concentrated in the 1960s and in the first half of the 1970s, and in only four cases with an induced reduction in disposable income above 1%. Unanticipated tax cuts are prevalent in the second part of the sample period, reaching 2% of disposable income in two cases (1972 and 1979).

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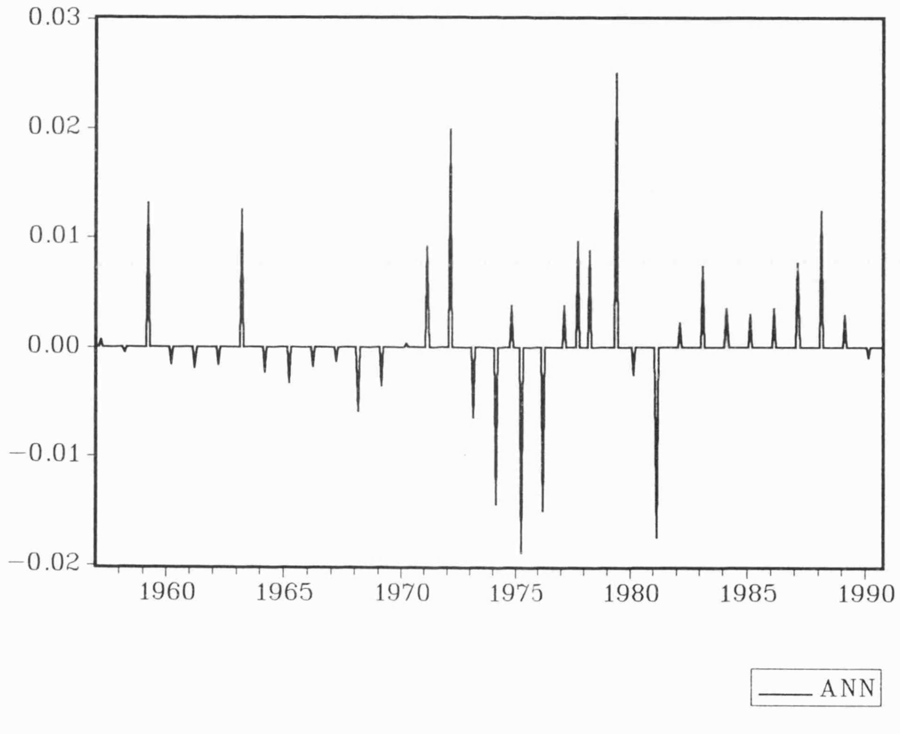
as his testing framework. Sumner allows for residual serial correlation in the estimation by means of generalized least squares procedures.

Figure 1

Unanticipated change in personal disposable income (as a fraction of disposable income) announced in Budget Statements 1960-1990.
(Positive values denote tax cuts)



The implementation of previously announced income tax changes is captured by a variable measuring the ratio of the estimated full year effect on tax revenues (from a *non-indexed* base) of proposed measures to personal disposable income, as a proxy for the percentage increase in personal disposable income *actually occurred* as a result of income tax changes. A simple "baseline" hypothesis has been made about the timing of implementation of such tax changes in order to capture the fact that, for institutional reasons, consumers' disposable income is affected by the tax measures announced in the Budget Statements only with a lag. We have assumed that the estimated effect on disposable income is uniformly distributed over the twelve months starting from the second to the fifth month after announcement in the Budget Report.[16] In the first month of implementation, beside the monthly quota of the tax change, adjustment for the period since the beginning of the tax year is made. For example, in the case of a tax cut yielding a 1% increase of disposable income announced in March but implemented with a two-month delay, we attribute a 2% increase in income (with respect to its level before the implementation) in the first month of implementation and a 1% increase afterwards. In terms of the monthly rate of growth of disposable income, this assumption implies that consumers faced an increase of 2% of their disposable income in the first month of the implementation, a reduction of 1% in the second month of implementation, and no change afterwards. The resulting variable measuring the monthly rate of change of disposable income, denoted *IMPL2*, is plotted in Figure 2 for the case of a two-month lag between the announcement and the implementation of income tax changes. Similar variables are constructed for three- to five-month lags and are denoted by *IMPL3*, *IMPL4* and *IMPL5* in the empirical analysis.[17]

[16] This assumption is consistent with the information contained in the Budget Reports and with the analysis of the 1976-1988 period in Sumner (1991).

[17] Some notes on particular episodes are in order. Most of the tax cuts promised in the April 1976 Budget were made dependent on *TUC* agreement on a low pay norm of "around 3 per cent". Such agreement was subsequently reached on the 5th of May, one month after the Budget date. As far as our test of the announcement effect is concerned, we attributed the whole estimated effect (£1224m), including the "conditional" £290m, to the Budget announcement month. Again in March 1977, part of the promised tax cuts -with an estimated effect on tax revenue of some £960m- were made contingent on negotiation of a new pay policy. Following the outcome of negotiation, in July only half of the originally announced tax cuts were implemented. We assumed the effect of such cuts to be £480m and added this figure to the £1303m of unconditional cuts announced in the March Budget (this amounts to assuming that people correctly anticipated the outcome of the pay negotiations and the subsequent Government response to it).

Figure 2

Implementation of income tax changes (as a fraction of disposable income) with one-month rebate payment 1960-1990.

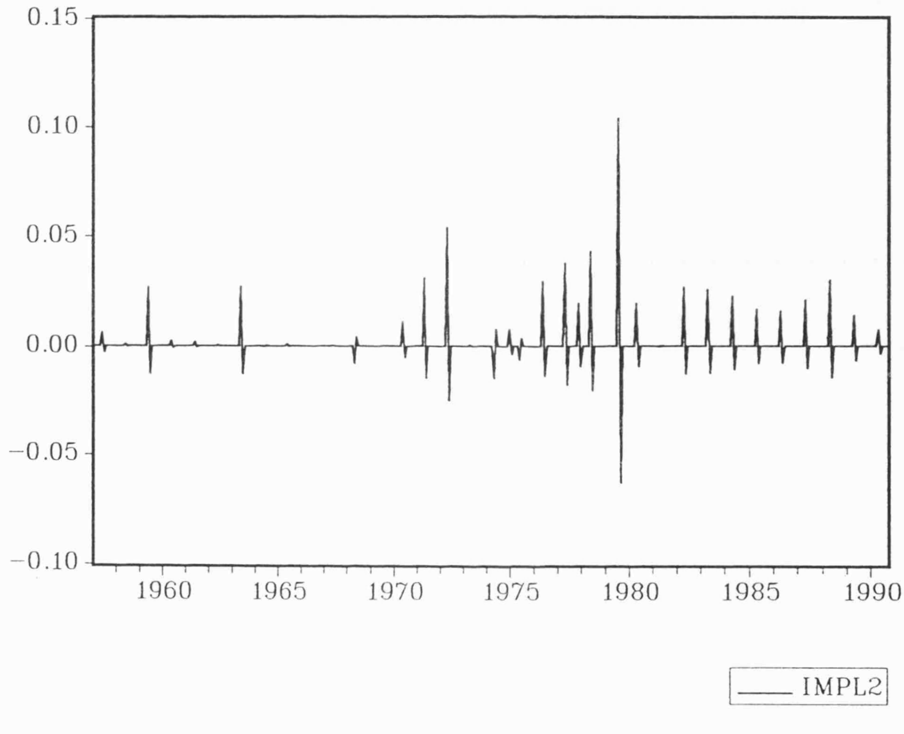


Table 2 reports sample correlations between the *ANN* and *IMPL* variables and a set of macroeconomic quantities representing various aspects of the business cycle. The very low values of these correlations (in only one case slightly higher than 0.10) indicates that our measures of fiscally-induced changes in disposable income can hardly be considered as proxies for other effects related to business cycle fluctuations.

Table 2
Correlation coefficients: 1960(1)-1990(9)

	ANN	*IMPL2*	*IMPL3*	*IMPL4*	*IMPL5*
Unemployment rate (Δ)	-0.06	-0.08	-0.05	-0.04	0.01
Inflation	-0.12	0.00	-0.01	-0.04	0.01
Interest rate (Δ)	-0.04	0.01	0.03	-0.04	0.01
Real earnings growth	0.08	-0.03	-0.01	0.06	-0.07
Share prices growth	-0.01	0.01	0.01	0.03	-0.01

Notes: The construction of the *ANN* and *IMPL* variables is described in the text. The other variables are: the monthly change in the unemployment rate, the monthly rate of change in the Retail price inde for All items, the monthly change in the nominal interest rate on 3-month prime bank bills, the monthly rate of change in average earnings in all industries (deflated with the Retail price index for all items), and the monthly rate of change of the Financial Times industrial share index (expressed in real terms, using the Retail price index for all items). For the share price variable the sample period is 1960(11)-1990(9).

To conduct our tests, the variables measuring the unanticipated announcement (*ANN*) and the delayed implementation of income tax changes (*IMPL*) are included in the Euler equation (8), together with monthly dummy variables (*M*), monthly trending seasonals (*MT*), and the expected real interest rate. The resulting equation, in its basic version, is:

$$\Delta c_t = \alpha + \sum_{i=1}^{11} \beta_{1i} M_{it} + \sum_{j=1}^{12} \beta_{2j} MT_{jt} + \gamma_1 ANN_t + \sum_{k=2}^{5} \gamma_k IMPLk_t + \phi E_{t-1} r_t + \varepsilon_t \qquad (9)$$

We are interested in testing the *REPI* hypothesis $\gamma_k=0$ $(k=2,\ldots,5)$ against the alternative that consumers react to the delayed implementation of pre-announced income tax changes, implying $\gamma_k>0$ $(k=2,\ldots,5)$. The strong evidence of an *MA(1)* error term in the estimated Euler equations reported above, implies that ordinary least squares estimates of the coefficient standard errors of (9) will not be consistent. Some transformation of the variables is needed in order to remove serial correlation in ϵ. The conventional backward (generalized least squares) transformation (as for example applied by Sumner (1991)), is not appropriate in the present context, since it may induce correlation between the transformed disturbances and the transformed regressors, leading to inconsistent estimates (Hayashi and Sims (1983) and Holden and Peel (1985)). Instead, we adopt the estimation technique suggested by Hayashi and Sims (1983), based on a "forward filtering" of the variables, followed by an instrumental variables estimation of the equation, using the untransformed regressors (and possibly lags thereof) as instruments. The forward filter is constructed by fitting an autoregressive model to the residuals (u) obtained from a first-stage estimate of (9).[18] Denoting by $g(L)$ the estimated polynomial in the lag operator such that $g(L)u_t=e_t$, and by y and x respectively the original dependent variable and the vector of regressors in (9), the suggested transformation of the variables is obtained by applying the coefficients in $g(L)$ to current and future values of y and x. The resulting transformed variables are: $y_t^*=g(L^{-1})y_t$ and $x_t^*=g(L^{-1})x_t$. When applied to (9) this transformation produces a serially uncorrelated disturbance which is a linear combination of current and future values of the original error term. Consistent estimates of the coefficients in (9) and their standard errors are then obtained by an IV regression of y^* onto x_t^*, using current and lagged values of x as instruments.

The presence of some residual serial correlation of order higher than one implies that, for some of our series, the current and some lagged values of the untransformed *IMPL* regressors (containing information known to agents at time *t-2* or earlier) may not be appropriate as instruments in the final instrumental variables regression. However, the particular time series behaviour of these regressors, shown in Figure 2, makes distant lags of the untransformed series unsuitable as instruments, being poorly correlated with the (transformed) variables that are to be instrumented. For this reason, in the following analysis, we nevertheless use the current and the first two lags of the untransformed *IMPL* regressors in the IV estimation, but provide a formal test for the adequacy of the instrument set. Only the current untransformed value is used as instrument for the *ANN*, *M* and *MT* regressors. The expected real rate is substituted by the actual rate (obtained by deflating the nominal rate, net of the standard tax rate, by the rate of change in retail prices) and instrumented with its own values at lags 3, 6 and 12, being significant regressors in a general 12th-order autoregression. Also the first lag was significant, but it is not a valid instrument here, due to the presence of an *MA(1)* disturbance term.

[18] The order of the autoregression for u is chosen with reference to the nature of the serial correlation in (9) and to the number of available observations in the sample (around 400 in our case). In the following empirical analysis we adopt an *AR(20)* model for the residuals u and explicitly test for the effectiveness of the procedure in removing residual serial correlation.

4. *Empirical results.*

The main results of our analysis are reported in Tables 3 to 5. We distinguish between a *basic* specification and an *augmented* specification of the consumption equations, the latter controlling for unanticipated news about agents' real income prospects other than fiscal innovations and for movements in the relative prices of sub-categories of goods. In addition to the set of seasonal dummy variables included in all equations using unadjusted data, a +1/-1 dummy variable is added in 1979(6)-1979(7) to control for the effect of a rise in the Value Added Tax, announced in the Budget Statement and implemented with a short delay. Since the resulting increase in consumption expenditure in anticipation of the future change in prices could be erroneously attributed to our announcement variable, intended to capture only the effect of announced income tax changes, it seems correct to separately control for this event. In the Food expenditure equation this variable was not statistically significant and is therefore excluded from the final specification.

Two types of test on the overall performance of the equations and on the adequacy of our estimation technique are presented for all regressions. The first is the Box-Pierce Q statistic for residual serial correlation, providing a check on the effectiveness of the Hayashi-Sims forward filtering in removing serial correlation up to the 24th order. The second is the Sargan test of overidentifying restrictions, constructed here as a Lagrange multiplier test from the regression of the estimated residuals of each equation on the whole set of instruments used. The resulting statistic, distributed as a χ^2 with degrees of freedom given by the number of additional instruments, provides a test for the validity of the instruments used in estimation.

The results for the basic specification are shown in Table 3. Looking first at the equation for expenditure on All Items (unadjusted data), we note that, with the exception of *IMPL3*, the variables measuring the implementation effect have positive coefficients, with a statistically significant magnitude at the fourth and fifth months after the Budget announcement (the point estimate of the elasticity of consumption expenditure to fiscally-induced changes in disposable income are 0.16 and 0.09 respectively), whereas the coefficient measuring the announcement effect is very close to zero. As for all other equations, the Q statistic confirms that serial correlation is removed and the Sargan test cannot reject the hypothesis of validity of the instrument set used in estimation.[19]

[19] Positive point estimates of the implementation effect, ranging from 0.05 to 0.14, are also found when seasonally adjusted data are used, with *IMPL2* having a statistically significant coefficient. The results obtained using seasonally adjusted data for expenditure on All Items are reported for

Table 3
Announcement and Implementation effects: basic specification
(Standard errors in parentheses)

Monthly rate of change of cons. exp. on:

Regressor	*Food*	*Clothing and Footwear*	*Household durables*	*All items*	*All items (adj. data)*
ANN_t	0.007	0.349	-0.172	0.031	0.149
	(0.256)	(0.678)	(0.703)	(0.304)	(0.226)
$IMPL2_t$	0.048	0.645**	0.331	0.168	0.141*
	(0.106)	(0.272)	(0.271)	(0.124)	(0.080)
$IMPL3_t$	-0.042	-0.013	0.197	-0.013	0.051
	(0.067)	(0.166)	(0.181)	(0.079)	(0.053)
$IMPL4_t$	-0.007	0.248*	0.185	0.156**	0.063
	(0.050)	(0.121)	(0.136)	(0.057)	(0.039)
$IMPL5_t$	-0.062	0.235**	0.074	0.087*	0.047
	(0.041)	(0.098)	(0.109)	(0.047)	(0.031)
$E_{t-1}r_t$	0.566	0.461	2.122	1.063	0.484
	(0.088)	(0.175)	(0.325)	(0.138)	(0.076)
$VAT79_t$	-	0.065	0.163	0.039	0.077
		(0.036)	(0.032)	(0.016)	(0.011)
R^2	0.95	0.94	0.86	0.98	0.48
σ *(x100)*	1.29	3.08	3.37	1.46	1.10
DW	2.02	2.07	2.04	2.09	2.05
Q(24)	9.3	14.1	4.2	19.9	19.5
Sargan (11)	11.8	16.3	15.1	18.1	11.7

Notes: Estimates obtained applying the Hayashi-Sims (1983) procedure described in the text. When seasonally unadjusted data are used a complete set of monthly dummy variables and of monthly trending seasonals are included in the equation. In the equation for Household Durables two additional *+1/-1* dummy variables are introduced in 1965(9)-1965(10) and 1975(4)-1975(5) to take care of outliers. The latter dummy is included also in the equation for All items (adjusted data). For the *IMPL* coefficients,** and * denote significance at the 1% and 5% level respectively. The *one-tail* critical values used are 1.64 (5%) and 2.33 (1%). σ is the standard error of the regression. *DW* is the Durbin-Watson statistic. *Q(24)* is the Box-Pierce statistic for residual serial correlation up to the 24th order, distributed as a $\chi^2(24)$; critical values are 32.2 (10%), 36.4 (5%), 43.0 (1%). *Sargan(11)* is the Sargan (1964) statistic, providing a test for the overidentifying restrictions and distributed as a $\chi^2(11)$; critical values are 17.3 (10%), 19.7 (5%), 24.7(1%). The sample period is 1960(1)-1988(12) (1960(11)-1988(12) for the All Items -adjusted data- series).

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completeness in the last column of all tables.

The results for our three sub-categories of goods show that expenditure on Clothing and Footwear strongly responds to the implementation of tax changes (with the statistically significant elasticity estimates ranging from 0.24 to 0.64), whereas expenditure on Food seems unaffected. For expenditure on Household Durables, the estimated coefficients on the implementation variables are all positive and display a more uniform pattern, but none of them is statistically significant. The announcement effect of unanticipated tax changes is statistically not well determined, with high standard errors on the relevant coefficients and high point estimates, if compared with those measuring the implementation effect, in the equation for Clothing and Footwear. The expected real interest rate enters all equations with highly significant coefficients. The point estimates indicate that Household Durable goods are the most sensitive to the real rate, whereas retail sales on all items display a unit elasticity of intertemporal substitution.[20]

In order to investigate whether some other news on agents' real income prospects, contemporaneous to the implementation of income tax changes, are at least partly responsible for the sizeable effect on consumption expenditure detected for some series, we augmented the basic specification presented in Table 3 with additional variables, capturing different types of news on consumers' income and wealth. To this aim, we first employed a measure of aggregate (all industries) real earnings. An estimate of the unanticipated movements in this variable (Δw) is obtained as the residual from a forecasting equation for the monthly rate of change in real earnings, including initially a complete set of monthly dummies and 12 lags of the dependent variable, the rate of change in industrial production, and the monthly change in the unemployment rate, and then reduced to include only statistically significant regressors. Using the same methodology, we constructed estimates of the unanticipated movements in real share prices and nominal interest rates. This latter variable has been suggested by Wilcox (1989) and Campbell and Mankiw (1991) as potentially relevant for consumption decisions on the ground that changes in nominal interest rates may have a direct influence on expenditure of indebted consumers who, facing an upper limit on the ratio of nominal debt service to nominal income, are forced to reduce consumption when nominal interest rates rise. Also Jackman and Sutton (1982), analyzing consumption decisions in the presence of imperfect capital markets, argued that increases in nominal interest rates, caused by unexpected increases in inflation, may have important effects on consumption levels of

[20] Removing the *VAT79* variable from estimation resulted, as expected, in an increase of the point estimate of the announcement effect (e.g. in the Clothing and Footwear equation the estimate yielded 0.55, with a standard error of 0.70). In all cases the point estimates remained not statistically significant and the results on the implementation variables were unaffected.

liquidity constrained individuals if, as is typically the case in the U.K., credit limits are not indexed.

Finally, we tried to capture relative price movements among the good categories considered by constructing estimates of the unanticipated rate of change of the three relative prices formed using the price indices of our individual retail sales series. These variables were obtained as residuals from autoregressions, specified starting from general 12th-order formulations, then reduced so as to include only significant regressors. These estimates, in addition to measuring the direct substitutability between food, clothing and household durables, may also be viewed as proxies for more general relative price movements between goods with a different degree of durability.

All "surprise" terms so constructed were included in a general augmented regression. The unanticipated changes in nominal interest rates and in real share prices yielded in-significant coefficients in all equations (with *t*-statistics always lower than *1*) and therefore were omitted in our final specifications. As for the relative price variables, only those concerning Household Durables and Clothing and Footwear *versus* Food (denoted respectively as $\Delta P(H/F)$ and $\Delta P(C/F)$) were statistically significant in at least one equation and consequently were retained, together with the unexpected change in real earnings, in the augmented equations shown in Table 4.[21] Due to the generated regressors problem (Pagan (1984, 1986)), the standard errors for the coefficients of the *ANN, IMPL* and $E_{t-1}r_t$ variables are derived from an IV regression which omits the surprise terms and includes, among the instruments, the variables used in the forecasting equations reported in the *Appendix*, with the only exception of those dated *t-1*, being not valid instruments in the present context.

The results previously obtained for the implementation variables are now confirmed for the Clothing and Footwear series and even strengthened for the Household Durables equation (now displaying two significant implementation coefficients, with point estimates of 0.23 and 0.38) and for the All items equation (where three implementation coefficients are significant with point estimates ranging from 0.11 to 0.20). The included surprise terms seem to affect somewhat the coefficient on the expected real rate variable, which decreases in all equations. Also some of the coefficients on the announcement variables are affected (especially in the Food and Household Durables equations), but in all cases they are unprecisely determined and not statistically different from zero. The unanticipated real earnings growth has a sizeable effect on most series, showing an elasticity close to unity for

[21] The forecasting equations used to generate the "surprise" variables included in the augmented equations are reported in the *Appendix* (section B).

Table 4
Announcement and Implementation effects: augmented specifications
(Standard errors in parentheses)

Regressor	Food	Clothing and Footwear	Household durables	All items	All items (adj. data)
	Monthly rate of change of cons. exp. on:				
ANN_t	-0.196	0.254	0.115	0.036	0.186
	(0.249)	(0.686)	(0.668)	(0.295)	(0.223)
$IMPL2_t$	0.031	0.652**	0.323	0.207*	0.157*
	(0.102)	(0.267)	(0.257)	(0.122)	(0.078)
$IMPL3_t$	-0.031	0.031	0.376*	0.063	0.086*
	(0.066)	(0.166)	(0.176)	(0.077)	(0.052)
$IMPL4_t$	0.032	0.268*	0.235*	0.187**	0.079*
	(0.049)	(0.120)	(0.129)	(0.056)	(0.039)
$IMPL5_t$	-0.023	0.247**	0.084	0.111**	0.063*
	(0.040)	(0.098)	(0.103)	(0.046)	(0.031)
$E_{t-1}r_t$	0.334	0.343	1.338	0.789	0.389
	(0.087)	(0.171)	(0.292)	(0.136)	(0.074)
$VAT79_t$	-	0.067	0.141	0.041	0.075
		(0.035)	(0.032)	(0.015)	(0.011)
Δw_t	0.092	0.274	0.960	0.345	0.240
	(0.069)	(0.173)	(0.189)	(0.084)	(0.060)
$\Delta P(H/F)_t$	0.165	-	-1.004	-	-
	(0.103)		(0.211)		
$\Delta P(C/F)_t$	0.517	-0.004	-	-	-
	(0.121)	(0.240)			
R^2	0.95	0.94	0.87	0.98	0.53
σ (x100)	1.21	3.06	3.18	1.43	1.06
DW	2.08	2.14	1.97	2.08	2.02
Q(24)	10.0	17.5	4.7	9.8	16.6
Sargan(11)	11.7	15.2	12.9	16.7	7.1

Notes: See notes to Table 3. Δw, $\Delta P(H/F)$, and $\Delta P(C/F)$ are the estimated residuals from forecasting equations for the monthly rate of change of real earnings (all industries), the relative price index of Household Durables versus Food and the relative price index of Clothing & Footwear versus Food. The standard errors for the coefficients of the *ANN, IMPL* and $E_{t-1}r$ variables are derived from IV regressions without the surprise terms and including, among the instruments, the variables used in the forecasting equations, only omitting those dated *t-1*, being not valid instruments in the present context.

Household Durables and a response of 0.34 of the All items series.[22] Finally, the sign pattern on the surprises in relative prices suggests some direct substitutability between sub-categories of goods.[23]

We now consider some additional issues concerning the robustness of our results. One potential problem with these estimates is due to the presence of measurement error in our measure of the tax cut effect, since it is based on the Treasury forecast which may differ from the actual effect on consumers' disposable income. Although Reilly and Witt (1990) have recently documented that the effect on tax revenues of income tax changes are better predicted by the Treasury than those for other categories of taxes, we nevertheless extended our instrumental variables procedure to overcome this problem. We included in the set of instruments, for each implementation variable, a dummy variable assuming the value of 1 in the months of implementation of tax cuts (-1 if a tax increase occurred) and zero otherwise. The pattern of results of Tables 3 and 4 is not affected, confirming that measurement error is not a relevant problem here.

We also assessed the robustness of the above results to a different assumption concerning the extent to which tax changes announced in the Budget were unanticipated by consumers. We made the alternative assumption that no tax-base indexation was expected by consumers in the 1960s and 1970s and for this part of the period we used the Treasury estimate of the effect on tax revenue calculated *from a non-indexed base* in the construction of the announcement variable (see the *Appendix* for details on these figures). The results obtained using this series again confirm those reported in Tables 3 and 4 for the implementation effect (in terms both of the elasticity estimates and of their statistical significance) and in only the case of the Household Durables and All items series, the coefficients on the announcement variable were larger than in our previous estimates: 0.90 (standard error 0.68) and 0.28 (0.30) respectively, in the augmented specifications.

Finally, we assessed the potential misspecification of the Euler equation (8) due to the omission of a (time-varying) conditional variance term. In fact, as shown by Caballero (1990b), in the presence of precautionary-savings behaviour (and with the assumption of an exponential utility function), the resulting Euler equation for the rate of growth of

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[22] The marked responsiveness of household durables to Δw is not inconsistent with the (S,s) model of durable expenditure of Bar-Ilan and Blinder's (1988).

[23] Our relative price measures, beside capturing direct substitution effects, may also be proxies for more general nondurables/durables price movements. The coefficient estimates must therefore be interpreted with caution. However, the strong significance of some of these terms suggests that they are successful in capturing some important relative price effects and then useful in evaluating the robustness of our main finding on the implementation variables.

consumption is: $\Delta c_t = (\theta E_{t-1}\sigma^2_{\epsilon t}) + \epsilon_t$. The expected variance of the disturbance term now enters the equation with a coefficient (θ) dependent on the degree of consumers' risk aversion. The omission of this potentially relevant variance effect should not affect the results obtained for the announcement effect, since our *ANN* variable is by construction orthogonal to the expected variance term. On the contrary, the *IMPL* variables are in the agents' information set at time *t-2* or earlier, so that, in principle, the inclusion of $E_{t-1}\sigma^2_{\epsilon t}$ could affect -in a way which is difficult to predict- the results. We tried to investigate the potential importance of this effect by testing for residual autoregressive conditional heteroscedasticity (*ARCH*) in Euler equations which omit the announcement and implementation variables. If no *ARCH* is detected, the variance at issue is likely to be constant in the sample period; therefore the omission of the conditional variance effect should not influence the estimates of the implementation coefficients. We obtained the following results for an *ARCH(12)* test (distributed as a $\chi^2(12)$, with a 5% critical value of 21.0): 16.8, 10.8, 10.7 and 5.4 for the Food, Clothing and Footwear, Household Durables and All items equation respectively. No evidence of *ARCH* behaviour is found in the Euler equation disturbance term: this may support the view that a conditional variance effect may not be too important in our sample.

Overall, we have found strong evidence that consumption expenditure reacts to the implementation of pre-announced income tax changes, a finding inconsistent with the basic *REPI* theory. The presence of liquidity constraints and the myopic behaviour of consumers are widely regarded as two of the main explanations for this kind of evidence. In principle, since our sample period displays some episodes of tax increases as well as a series of tax cuts, it should be possible to discriminate between the above explanations by separating the effect of announcements and implementations of tax changes of different sign. In fact, if liquidity constraints prevent the increase in consumption following announcements of future income tax cuts, they do not prevent a downward adjustment when future income tax increases are announced. On the other hand, myopic behaviour would imply no response in both cases. As for the implementation effect, under liquidity constraints only the implementation of tax cuts should affect consumption (since downward adjustment of expenditure in the face of tax increases should have already taken place at the announcement date), whereas the implementation of tax changes of either sign should affect consumption under myopia. However, when we split the announcement and implementation variables in order to perform a simple test of liquidity constraints *versus* myopia, the results were not conclusive. More specifically, very high standard errors of the coefficient estimates did not allow reliable inferences for the announcement effect (although again none of the *ANN* variables was significant). As for the implementation variables, those capturing the effect of tax cuts replicated the pattern (in terms of point estimates and statistical significance) of the results previously presented, whereas those relating to tax increases yielded very unprecisely estimated coefficients (none of them statistically significant). One representative example, concerning the All items series, is reported in Table 5, where variables denoted by (-) and (+) refer to announcement and implementation of tax cuts and tax increases respectively, so that the (+) variables have negative values in the relevant months. The specification of the equation is that of table 4, including surprise terms. Although, taken literally, these results do not reject the hypothesis of liquidity constraints, the lack of precision of the estimates of some important coefficients prevent us from drawing any sharp inference about the two competing hypotheses. Perhaps the limited variability of the regressors capturing the implementation effect of tax increases (occurred only five times in the sample, in 1964, 1968, 1974, 1975 and 1981, and for small fractions of disposable income) is responsible for such unprecise estimation results.

Finally, two points concerning the economic interpretation of our results must be addressed.

First, we detected a positive effect of the implementation of pre-announced income tax changes on consumption expenditure on durable and semi-durable items (but at the monthly frequency the degree of durability of the latter goods is very high), whereas food consumption does not react to anticipated changes in disposable income. Although there is no well-developed theory of consumption capable to formally explain this fact, it does not seem too implausible to think that the delayed implementation of tax cuts, resulting in increases in disposable income having for a sizeable part the nature of a one-time rebate payment, maybe adding to previous savings, triggers the purchase of some clothing or household durable item, instead of determining an increase in consumers' food expenditure. On the empirical side, this result is not a peculiarity of UK data, since it is qualitatively similar to the findings of Poterba and Summers (1987) and Wilcox (1989) for the US.

Second, the numerical pattern of coefficients' estimates on the implementation variables indicates that most of the effect occurs in the second, fourth and fifth months after Budget announcements for expenditure on Clothing and Footwear and on All items, and in the third and fourth months for expenditure on Household Durables. The fact that, during the thirty years of our sample, the implementation of tax changes may have started with a variable delay with respect to the Budget statement (but mostly from the second to the fifth month after announcement) and continued gradually for several subsequent periods, makes it difficult to account for the implementation coefficient pattern with any simple hypothesis. In fact, such coefficients may capture both the impact effect on consumption of changes in

Table 5

Asymmetric response to announcement and implementation of income tax changes.

(Standard errors in parentheses)

Dependent variable:
monthly rate of change of consumption expenditure on All items

Announcement effect		*Implementation effect*	
$ANN(-)_t$	0.076		
	(0.431)	ANN_t	0.118
$ANN(+)_t$	-0.035		(0.310)
	(0.534)		
		$IMPL2(-)_t$	0.221*
$IMPL2_t$	0.242*		(0.128)
	(0.120)	$IMPL2(+)_t$	-0.429
			(0.958)
		$IMPL3(-)_t$	0.069
$IMPL3_t$	0.076		(0.083)
	(0.080)	$IMPL3(+)_t$	-0.387
			(0.583)
		$IMPL4(-)_t$	0.191**
$IMPL4_t$	0.184**		(0.059)
	(0.057)	$IMPL4(+)_t$	0.188
			(0.418)
		$IMPL5(-)_t$	0.128**
$IMPL5_t$	0.114**		(0.049)
	(0.046)	$IMPL5(+)_t$	-0.581
			(0.372)
$E_{t-1}r_t$	0.842	$E_{t-1}r_t$	0.837
	(0.148)		(0.145)
Δw_t	0.350	Δw_t	0.363
	(0.084)		(0.084)
$VAT79_t$	0.045	$VAT79_t$	0.044
	(0.016)		(0.017)
R^2	= 0.98	R^2	= 0.98
$\sigma(x100)$	= 1.44	$\sigma(x100)$	= 1.46
DW	= 2.20	DW	= 2.23
$Q(24)$	= 13.1	$Q(24)$	= 14.5
$Sargan(11)$	= 16.5	$Sargan(11)$	= 17.1

Note: see the notes to Table 3. Variables denoted by (-) and (+) refer to announcement and implementation of tax cuts and tax increases respectively, so that the (+) variables have negative values in the relevant months. In the announcement effect equation the overidentifying instruments are those employed in Tables 3 and 4. In the implementation effect equation we used one lag for each of the eight implementation variables and three lags for the interest rate variable.

income occurred in each particular month and -with the exception of the coefficient on the first *IMPL* variable- the effect of a slow adjustment of consumption expenditure to income variations, an hypothesis which may apply, at the monthly level, not only to Household Durables, but also to Clothing and Footwear goods.

Moreover, even though the estimates of the implementation effects are sufficiently precise to strongly reject the null hypothesis tested, the differences among the various coefficients are not statistically significant. In order to show this, we formally tested a simple baseline hypothesis, imposing equality of the four implementation coefficients in the augmented specifications of Table 4. Here we report the value of the likelihood ratio statistic obtained (distributed as a χ^2 with three degrees of freedom, with 6.25 as the 10% critical value) and the estimate (and standard error) of the unique constrained implementation coefficients, providing a summary measure of the response of expenditure on various goods to income changes:

	Clothing and Footwear	*Household Durables*	*All items*	*All items (adj. data)*
LR(3)	3.56	3.83	5.10	2.43
IMPL	0.233^{**}	0.171^{*}	0.127^{**}	0.079^{**}
	(0.072)	(0.088)	(0.037)	(0.025)

The hypothesis of equality of all implementation coefficients is clearly not rejected and the coefficient estimates confirm elasticities of expenditure on durable items (0.23 and 0.17) higher than the overall response of consumption (0.13). Given the above results, we believe that the apparently irregular pattern of coefficients presented in Tables 3 and 4 does not affect the interpretation of our main finding as a strong response of consumption expenditure to anticipated changes in current income.

5. *Conclusions.*

The response of aggregate consumption to current income fluctuations has always been the focus of the empirical evaluation of the rational expectations permanent income model of consumption. According to the *REPI* model, pre-announced income tax changes, determining variations in consumers' disposable income, should not have any effect on current consumption, their effect being already included in consumption levels at the time of the announcement. Our extensive analysis of a long series of such episodes for the U.K. provides strong evidence against the basic version of the *REPI* model (and the Ricardian Equivalence proposition). In fact, consumption expenditure positively reacts to fiscally-induced movements in disposable income only at the implementation date. The overall effect is clearly attributable to the semi-durable and durable components of consumption expenditure.

Appendix

A. *Construction of a measure of the unanticipated part of income tax changes announced at Budget dates.*

For the final part of the sample period (1982-1990) the FSBR reports separate figures for the estimated change in income tax revenue both from an *indexed* and a *non-indexed* base. Assuming that agents correctly predicted the extent of indexation decided by the authorities (which was not uniform, despite the Rooker-Wise amendment establishing allowance indexation as the rule since 1977 (Sumner (1991)) and did not foresee any discretionary change in income taxation, we used the change in tax revenue from an *indexed* base as our measure of the unexpected tax change announced in the Budget Statements. For the previous period only the Treasury estimate calculated from a *non-indexed* base is available and constructing a proxy for expected indexation is not straightforward. Moreover, it could be argued that if, at least in the 1960s and early 1970s, consumers were slow to recognize the existence of inflation and governments did not immediately respond to it with tax-base indexation, the use of the available Treasury estimate -with no correction- for the pre-1981 period may be justified (we owe this point to M. Sumner). We adopt this strategy in the results section, when checking the robustness of our findings. Here, we construct a rough proxy for the unexpected part of the announcement, using the following equation:

$$[\Delta Tax\ Rev.(indexed)]_t = [\Delta Tax\ Rev.(non\text{-}indexed)]_t - k \cdot \pi_{t-1} \cdot [Total\ Tax\ Rev.]_t$$

Our estimate of the change in income tax revenue -after allowing for base-indexation- resulting from the Budget for year t is equal to the Treasury estimate of the change in income tax revenue calculated from a non-indexed base minus a term correcting for expected indexation to the inflation rate in year $t-1$ (π_{t-1}). The coefficient k, ranging from zero (no expected indexation) to one (in the case of complete indexation), is obtained from the estimation of the above equation over the period 1982-1990, when both the indexed and non-indexed Treasury estimates of tax revenue change are available, and is set equal to *0.523*. Then, according to our proxy, income tax changes announced in the Budget Statements during the 1960-1981 period were expected to reflect nominal allowance indexation to around half of the inflation rate occurred in the year preceding the Budget. We chose, as the relevant measure of past inflation, the rate of change of the GNP deflator calculated for the year ending in the December preceding Budget announcements. In other words, we estimated for the 1980s the part of the difference between the two Treasury estimates of tax revenue change that can be related to past inflation (which is in the consumers' information set at the dates of Budget announcements) and used this estimate (k) for the whole sample.

Finally, for the whole period, the variable so constructed has been divided by twelve and expressed (after changing sign) as a ratio to personal disposable income to obtain a measure of the monthly percentage (unexpected) variation in disposable income perceived at Budget announcements.

B. *Forecasting equations used to generate "surprise" terms.*

The unanticipated changes in the growth rate of real earnings and in the relative price level of different categories of goods are constructed as residuals from the following forecasting equations for ΔW, $\Delta P(H/F)$ and $\Delta P(C/F)$:

$$
\begin{aligned}
\Delta W_t = {} & \underset{(0.055)}{-0.236}\Delta W_{t-1} \underset{(0.056)}{-0.095}\Delta W_{t-2} \underset{(0.056)}{+0.158}\Delta W_{t-4} \underset{(0.054)}{-0.039}\Delta W_{t-6} \underset{(0.052)}{-0.118}\Delta W_{t-7} \\
& \underset{(0.053)}{-0.099}\Delta W_{t-10} \underset{(0.054)}{-0.117}\Delta W_{t-11} \underset{(0.025)}{-0.030}\Delta Y_{t-1} \underset{(0.025)}{-0.047}\Delta Y_{t-2} \underset{(0.024)}{-0.068}\Delta Y_{t-4} \\
& \underset{(0.022)}{-0.040}\Delta Y_{t-12} \underset{(0.004)}{-0.008}\Delta U_{t-2} \underset{(0.004)}{-0.006}\Delta U_{t-8} \underset{(0.004)}{+0.008}\Delta U_{t-10} + \mathit{seas.}
\end{aligned}
$$

$R^2=0.53 \qquad \sigma=1.05\% \qquad Q(24)=20.2$

$$
\begin{aligned}
\Delta P(H/F)_t = {} & \underset{(0.054)}{0.138}\Delta P(H/F)_{t-1} \underset{(0.054)}{-0.118}\Delta P(H/F)_{t-2} \underset{(0.053)}{-0.139}\Delta P(H/F)_{t-4} \underset{(0.054)}{+0.071}\Delta P(H/F)_{t-5} \\
& \underset{(0.053)}{-0.108}\Delta P(H/F)_{t-6} \underset{(0.053)}{+0.100}\Delta P(H/F)_{t-7} \underset{(0.053)}{+0.088}\Delta P(H/F)_{t-9} \underset{(0.053)}{+0.077}\Delta P(H/F)_{t-10} \\
& \underset{(0.053)}{-0.082}\Delta P(H/F)_{t-11} + \mathit{seas.}
\end{aligned}
$$

$R^2=0.43 \qquad \sigma=0.87\% \qquad Q(24)=28.8$

$$
\begin{aligned}
\Delta P(C/F)_t = {} & \underset{(0.055)}{0.115}\Delta P(C/F)_{t-1} \underset{(0.055)}{-0.078}\Delta P(C/F)_{t-4} \underset{(0.055)}{+0.081}\Delta P(C/F)_{t-7} \underset{(0.054)}{+0.061}\Delta P(C/F)_{t-8} \\
& + \mathit{seas.}
\end{aligned}
$$

$R^2=0.49 \qquad \sigma=0.83\% \qquad Q(24)=10.6$

In all equations a complete set of monthly dummy variables and monthly trending seasonals is included. The sample period is 1960(1)-1990(9).

C. *Variable description and data sources.*

(i) To construct the series for *real Retail Sales*, the following variables have been used (January 1980=100):

Variable Description	*Source*
Retail Sales Value indices (Food, Clothing & Footwear, Household Durables and All Items) *not seas. adjusted*	*MDS*
Retail Sales Volume index *seas. adjusted*	*MDS* (1960-1962) *ET* (1963-1990)
Retail Price indices (Food, Clothing & Footwear, Household Durables and All Items) *not seas. adjusted*	*RPI* (1960-1973) *MDS* (1974-1990)

(ii) The *Tax Change* measure was constructed using the following variables:

Estimated full-year effect of changes in inc. taxation	*FSBR*
Personal disposable income (quarterly figure)	*NA*

(iii) Other variables used were:

Y	(Log of) Index of Production (all industries)	*MSD*
S	(Log of) Financial Times industrial share index	*FS*
R	Interest rate % 3-month prime bank bills	*ET*
U	Unemployment rate	*MDS*
W	(Log of) Average real earnings index (all industries) deflated with the Retail Price index (all items)	*MDS*

ET: Central Statistical Office, *Economic Trends* (various issues)
FS: Central Statistical Office, *Financial Statistics* (various issues)
FSBR: *Financial Statement and Budget Report*
MDS: Central Statistical Office, *Monthly Digest of Statistics* (various issues)
NA: Central Statistical Office, *National Accounts* (1991)
RPI: Department of Employment, *Retail Price Indices, 1914-1986* (1987)

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Rome, 30 October 1995

DECLARATION

I hereby certify that the following chapter of the thesis

Money, Policy Regimes, and economic fluctuations
by
Fabio-Cesare Bagliano

describes cojoint work with the undersigned:
Chapter 4 Active monetary policy and the real effects of nominal shocks. A comparison of the interwar and postwar U.S. experience.

The contribution of Fabio-Cesare Bagliano to the above chapter has been at least 50%.

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DECLARATION

I hereby certify that the following chapters of the thesis: "Money, policy regimes and economic fluctuations" by Fabio-Cesare Bagliano describe conjoint work with the undersigned:

Chapter 2 Money Demand Instability, expectations and policy regimes: an application to Italy (1964-1986)

Chapter 5 Monetary policy, credit shocks and the channels of monetary transmission. The case of Italy 1982-1994.

The contribution of Fabio-Cesare Bagliano to each venture has been of at least fifty per cent.

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