BIS WORKING PAPERS
No. 18 - April 1992

EUROPEAN INTEGRATION AND THE DEMAND FOR BROAD MONEY

by C. Monticelli and M-O. Strauss-Kahn

BANK FOR INTERNATIONAL SETTLEMENTS
Monetary and Economic Department

BASLE

December 1991

EUROPEAN INTEGRATION AND THE DEMAND FOR BROAD MONEY

Carlo Monticelli and Marc-Olivier Strauss-Kahn*

ABSTRACT

Co-integration analysis and error-correction modelling show that a stable and predictable aggregate demand for broad money holds for the group of countries participating in the Exchange Rate Mechanism (ERM) of the European Monetary System. This result seems robust to different econometric techniques, samples and methods of conversion of national variables into ecus. Furthermore, ERM-wide equations compare favourably with national equations and with the previous estimates of area-wide equations for narrow money. These findings suggest that an area-wide broad monetary aggregate can play a useful role in the co-ordination of monetary policy at the European level.

---

* Economic Unit of the Committee of Governors of EC Central Banks.
The authors gratefully acknowledge comments by their colleagues and by economists from Community central banks, in particular J. Dolado (Banco de España) and J. Henry (Banque de France). C. Boersch and H. Christiansen have provided valuable research assistance.

EUROPEAN INTEGRATION AND THE DEMAND FOR BROAD MONEY

Table of contents

EUROPEAN INTEGRATION AND THE DEMAND FOR BROAD MONEY

INTRODUCTION

In recent years, European economies have experienced an increasing degree of financial integration. Administrative barriers to capital movements have been progressively removed while the Exchange Rate Mechanism (ERM) of the European Monetary System (EMS) has brought about greater stability in the exchange rates of participating countries. These developments raise the issue of whether a stable aggregate money demand function can be estimated for the countries participating in the ERM, as an ERM-wide monetary aggregate would contain useful information about the economic conditions and the monetary stance of the area as a whole. Such an aggregate would thus be relevant to the conduct of monetary policy at the European level, either through the co-ordination of national monetary policies or, prospectively, through the management of a single policy by the European System of Central Banks. The empirical evaluation of the properties of an ERM-wide demand for money may also supply qualitative indications on the current extent of European financial integration, providing insights into the proximate causes underlying the volatility in money demand experienced in various European countries.

This paper draws on recent developments in applied money demand analysis, especially those related to co-integration and error-correction models (ECM),[1] and presents several novel features as compared to the other studies which estimated area-wide aggregate money demand functions for some European countries (Bekx and Tullio, 1989; Kremers and Lane, 1990). Firstly, broad money stocks are considered, while previous work focused on narrow definitions of money. Broad aggregates, rather than narrow ones, are currently used as targets or indicators in a majority of European countries and hence they may be expected to have a greater bearing on the implementation of monetary policy. Secondly, the aggregation involves all ERM countries, including Spain and the United Kingdom, whose currencies

---

1 The classic reference papers are Hendry (1980, 1985) on ECM, and Granger (1983), Engle and Granger (1987), and Johansen (1988) on co-integration.

are currently in the wide band of the ERM.[2] Finally, the dynamics and the robustness of the equations for money demand are examined by applying different econometric techniques and by repeating the analysis for two different samples and for two different methods of conversion of national variables into the same currency (ECU).

The econometric evidence in this paper supports the existence of a stable relationship between broad money, income and the interest rate (and variations in the dollar/ECU exchange rate), which has held for the ERM area at least since the start of the EMS in 1979. The degree of economic and financial integration within the Community seems sufficiently high for an ERM-wide demand for broad money to compare well with single-country equations, outperforming some of them in terms of stability or predictability.

The paper is divided into five sections: Section I discusses the rationale for assessing the stability of money demand at the ERM level; Section II describes the econometric approach adopted; Section III deals with the definition of area-wide variables; Section IV comments on the empirical results; and Section V highlights the main conclusions.

I. RATIONALE FOR ASSESSING AN ERM-WIDE MONEY DEMAND

In the 1970s, the setting of monetary policy in several countries hinged on the targeting of some measure of money whose demand was perceived to be stable. In contrast, in the 1980s, financial innovation and deregulation had a destabilising impact on the relationship between money and the final variables, leading many central banks to adopt a more eclectic approach based on various indicators. Partly as a result of the development of new econometric techniques, confidence in the stability of money demand has revived in recent years. This has contributed to maintaining (or restoring) the role of monetary aggregates as targets or indicators in the conduct of monetary policy within the Community. In some countries, aggregates are viewed as providing the intermediate link between

---

2 However, Luxembourg, which forms a monetary union with Belgium, is omitted for statistical reasons, as in previous studies. The exclusion of Greece and Portugal (not yet participating in the ERM in 1991) implies an underestimation of the Community of around 3% in terms of GDP or money stocks.

changes in monetary instruments and developments in nominal income, while in others they are considered mainly as timely indicators of developments in nominal income, which would otherwise be known only with additional lags. In both cases, the closer and the more stable the relationship between money and the final variables, the more predictable is the path of monetary aggregates on the basis of current information, and the greater is the confidence which can be attached to the signal money gives about the developments in nominal aggregate demand. These arguments account for the considerable attention paid to the modelling of money demand equations. The most up-to-date national results on broad money show that a stable demand equation holds from the 1970s to the second half of the 1980s in the majority of cases, but not always with satisfactory dynamic properties and forecasting performances.[3]

The difficulties encountered in the specification of stable, well-performing single-country equations supply the first rationale for studying money demand at the ERM level. They suggest that, besides deregulation and innovation at the national level, the growing degree of international integration also has a major impact on national money demands. As international integration progresses, residents' demand for monetary assets is increasingly influenced by the rates of return on foreign assets, whether monetary or not. Thus, foreign variables - such as exchange and foreign interest rates - become more relevant for estimating money demand functions, although in practice they are only rarely included in single-country equations. This is partly due to the well-known technical problems associated with the high degree of correlation between national and foreign interest rates (correlation which becomes stronger the higher the degree of asset substitutability) and partly to the difficulties in identifying the relevant foreign variables. These complications are reduced when money demand is estimated at the area level. Area-wide equations help

---

3 The following papers were used as reference: Johansen and Juselius (1990) for Denmark; Schmid and Herrman (1991) for Germany; Dolado and Escrivà (1991) for Spain; De Bandt (1991) for France; Hurley and Guiomard (1990) for Ireland; Muscatelli and Papi (1990) for Italy; Fase and Winder (1991) for the Netherlands; Hall et al. (1990) for the United Kingdom. For recent cross-country studies, see Boughton (1991a, b) and Boughton and Tavlas (1991).

overcome the specification problems with respect to "within-the-area" foreign variables, while rendering the influence of third countries' variables sharper and hence arguably easier to capture.

However, an ERM-wide equation for money demand qualitatively differs from a mere averaging of single-country demands also for other reasons. Economic and financial integration increases the size and importance of cross-country spillover effects. This entails a stronger direct impact on national money demands of quantity variables from other ERM countries (such as income and wealth), which adds to their indirect effect through interest and exchange rates as well as through domestic quantities. As a result, shocks in money demand become more likely to be correlated across countries. Furthermore, the exchange rate stability ensured by the ERM, together with the progress towards Economic and Monetary Union (EMU), enhances the scope for currency substitution within the area. Besides the use of foreign currency to carry out domestic transactions (currency substitution stricto sensu), exchange rate stability fosters the holding of transaction balances in foreign currencies for spending abroad and, more generally, the holding of (monetary) assets denominated in foreign currency. Currency substitution thus gives rise to further reasons for the strengthening of interlinkages between national money demands, possibly increasing their volatility when considered separately. By internalising the cross-country spillover effects within the Community, an area-wide money demand provides a fruitful device to tackle the above issues, although at the cost of imposing the restriction that money demand functions have the same structure in all countries. The encouraging results obtained from the empirical analysis on ERM-aggregate variables suggest that this price is worth paying.

Another rationale for assessing the properties of an ERM-wide money demand comes from its bearing on the conduct of monetary policy at the European level. The importance of international financial linkages in the environment of free capital mobility and exchange rate stability which characterises Stage One of EMU is weakening the degree of controllability of the single country's money stock. This requires the co-ordination of monetary policies aimed at approximating a "common" policy - or the institutional framework necessary for the implementation of a single monetary policy in Stage Three - in order to regain the controllability of the money supply at the area level, under a regime of flexible exchange rates vis-à-vis third currencies. In this respect, the assessment of the stability of the ERM-wide money demand plays an important role, since a stable relationship between monetary aggregates and final variables is a necessary, although not sufficient, condition for their usefulness in the setting of monetary policy. Furthermore, an experimentation with the money stock which results from adding up national aggregates as currently defined can provide a useful benchmark to assist in the definition of alternative area-wide money concepts in terms of asset composition, sectoral definition and application of the criteria of residence of the holder, location of the issuer and currency denomination (Goodhart, 1989).

II. CO-INTEGRATION AND ERROR-CORRECTION STRIKE AGAIN

Applied research has long since come to acknowledge that the relationship postulated by money demand theory may hold in the "long-run, steady-state equilibrium", but that other factors, such as adjustment costs and financial innovation, concur to determine the short-run dynamics of money balances. Co-integration and error-correction modelling have addressed this issue afresh and the related spate of methodological literature has produced a wide range of different techniques and tests. Rather than focusing on a particular approach, this paper employs different econometric procedures to assess the robustness of the empirical findings. The steps followed in the empirical analysis, summarised in Table 1, are presented briefly below.

1. Stationarity and co-integration

As a starting point, consider the reduced form of the demand for nominal money [M]. It may be expressed as a function of the price level [P], a scale variable in real terms [Y] and a set of other variables [R(i)]. The latter may capture the return on the components of M and the returns on alternative assets, typically domestic short and/or long-term financial assets. The long-run demand for money balances can be specified in real terms, assuming unitary elasticity of money to the price level (for a discussion see, for instance, Hendry, 1986, and Boughton, 1991a). Converting all variables, except rates, into logarithms (denoted by small letters), the long-run demand

TABLE 1: PROCEDURE

Tests	Object	Tests/Tables
STATIONARITY	Order of integration	DF/ADF : Dickey - Fuller, (1981) $t\alpha$(unit root), $t\mu$(cst.), $t\beta$(trend) Φi (i=1,2,3 : joint tests): Perron - Phillips, (1988)
COINTEGRATION	Existence of a cointegrating relationship Uniqueness of the cointegrating vector	DF/ADF : Engle - Yoo, (1987) ADF : Phillips - Ouliaris, (1990) Trace, Lambda max : Johansen (1988) Johansen - Juselius, (1990)
LONG-RUN PROPERTIES	Comparison between the long-run estimates in the first-stage OLS: - & the maximum-likelihood estimates from the Johansen method - or the long-run solutions of the single-stage ECM	 Chi² F-test
SHORT-RUN DYNAMICS	Specification adequacy and predictive power of the: - single-stage ECM - second-stage ECM	Usual diagnostic tests: LM(1-1); LM(1-5); LM(1-10) NORMALITY; ARCH 4 HET (Square Residuals) RESET Out-of-sample Hendry and Chow tests

for real balances, denoted by a star, can be written as follows:

$$[1] \qquad (m-p)^{*} = k + \mu\, y + \sum_{i} \beta(i)\, R(i)$$

Traditional econometric inference can be consistently applied only if the above variables are covariance stationary, i.e. integrated of order zero. Testing the order of integration of the variables is therefore a necessary preliminary step to the empirical analysis. To this end, several tests are available in the literature (see Dolado et al., 1990, for a survey). In the present analysis, the results of the DF/ADF tests (Dickey and Fuller, 1981) are compared to those of the tests proposed by Phillips and Perron (1988). As seasonally-adjusted time series were used, the issue of seasonal integration is not tackled (see Osborn et al. (1988) and Hylleberg et al. (1990) for a discussion).

If the time series appearing in equation [1] are non-stationary with stationary first differences, i.e. they have unit roots,[4] then a stable long-run relationship between (non-stationary) variables exists when co-integration occurs, i.e. when a linear combination of these variables is found to be stationary. Operationally, this amounts to testing the stationarity of the residual of the OLS static regression [1] and is performed comparing the DF/ADF test statistics with tables from both Engle and Yoo (1987) and Phillips and Ouliaris (1990). The empirical analysis also resorts to the maximum likelihood estimation method proposed by Johansen (1988) and Johansen and Juselius (1990), which considers explicitly the question of non-unique co-integrating vectors in a multivariate setting.[5]

---

4 For the sake of simplicity, the case of a subset of I(2) variables that co-integrate CI(2,1) is not considered here; see Engle and Granger (1987).

5 This method establishes the number of existing co-integrating relationships by carrying out a sequence of parametric tests that there are r or fewer co-integrating vectors ($0 \le r \le n-1$, where n is the number of variables in the model) both against a general alternative (trace test) or against the alternative of r+1 vectors (maximum eigenvalue test).

2. Long-run estimates

Once the existence of a long-run money demand function has been established, the issue arises of how to estimate its parameters. The super-consistency theorem (Stock, 1987)[6] supports the use of the OLS estimation of the static model. In contrast, other studies - for example, Banarjee *et al.* (1986) - warn against the problems involved with the finite sample bias, while Johansen (1988) points out that the OLS estimates are conditional on the arbitrary normalisation implicit in the selection of the dependent variable for the regression equation. To gain a better insight into the long-run estimates, the first-stage OLS estimates (from equation [1]) are compared with those obtained from the Johansen method and from the long-run solution of the single-stage ECM (from equation [2] below). The comparison is performed by carrying out parametric tests for the hypothesis that the long-run parameters are, individually and jointly, equal to the first-stage OLS estimates.

Testing the equality of estimates across methods has a noteworthy implication for the maximum-likelihood estimates. When k restrictions are imposed on the long-run parameters of the VAR system underlying the Johansen procedure, the maximum number of different co-integrating vectors is reduced by k. In testing for the joint equality of maximum-likelihood estimates with first-stage OLS ones, $k = n-1$ (where n is the number of variables of the system); hence, all normalised co-integrating vectors collapse into one. The non-rejection of the restrictions therefore provides evidence on the uniqueness of the co-integrating vector which is additional to that obtained from the trace and maximum eigenvalue tests.

3. Error-correction modelling

The final stage of the empirical analysis is the specification of the ECM under the assumption that actual real money balances adjust only progressively to desired real balances. The change in effective real balances $\Delta(m-p)$ may be expressed as an

---

6 The theorem shows that OLS estimates of the co-integrating vector converge to the true parameters at a rate proportional to the inverse of the sample size, instead of its square root as in the usual OLS case.

error-correction model:

$$[2] \qquad \Delta(m-p) = -c + [\, a(m-p)_{t-1} + b\, y_{t-1} + \sum_i d(i)\, R(i)_{t-1} \,] + [\, \sum_i e_j \Delta y_{t-j} + \sum_i \sum_j f(i)_j \,\Delta R(i)_{t-j} \,]$$

The first part of the right-hand side of equation [2] (lagged levels) represents the adjustment of actual to desired balances in the previous period, as defined by equation [1]. The second part of the right-hand side of equation [2] (lagged variations) represents changes in desired money balances.

The error-correction term may alternatively be written as $\hat{\varepsilon}$, the residual of equation [1] estimated on actual real money balances, (see Engle and Granger, 1987):

$$[1'] \qquad (m-p) = \hat{c} + \hat{\mu}\, y + \sum_i \hat{\beta}(i)\, R(i) + \hat{\varepsilon}$$

$$[2'] \qquad \Delta(m-p) = -c + a\, \hat{\varepsilon}_{-1} + \sum_j e_j\, \Delta y_{t-j} + \sum_i \sum_j f(i)_j\, \Delta R(i)_{t-j}$$

Provided the set of variables are co-integrated, the ECM may therefore be estimated by a single-stage (equation [2]) or a two-stage approach (equation [2']), as shown by the Granger Representation Theorem (Granger, 1983: Granger and Weiss, 1983). Both approaches are applied in the paper.

III. DEFINITION OF ERM-WIDE VARIABLES

The definition of area-wide variables involves two main issues: the selection of the national variables to be aggregated and the choice of the aggregation method.

1. Selection of national variables

As the names of broad money aggregates are often misleading (for example, Italian M3 is "broader" than U.K. M4), a careful comparison of national definitions is necessary to identify aggregates that are reasonably similar in terms of the types of asset they include. To improve their similarity, some changes were implemented in line with the new definitions announced for 1992 in some countries. They mainly consist in the generalisation to all ERM countries of the following characteristics: the inclusion of monetary assets denominated in foreign currency (in Spain and the United Kingdom); the exclusion of short-term Government liabilities (such as Treasury bills, in Belgium, Spain and the Netherlands) and of long-term non-negotiable certificates (in France and Italy). A detailed list of the aggregates selected and of the changes made is reported in Appendix 1.

The elimination of major inconsistencies in the asset composition of aggregates does not ensure, however, the avoidance of omissions or duplications in the area-wide monetary measures which result from adding up national aggregates (see Goodhart, 1989). As in previous studies, the empirical analysis of this paper underestimates the actual ERM-wide money stock, since it does not include cross-border holdings (CBHs) in monetary assets by European residents within the Community. The unreliability of the existing data and the short period for which they are available discourage any attempt to include CBHs in the area-wide broad aggregate at present.[7] Not only are these CBHs still of a fairly limited size, but national definitions extended to (some) CBHs, where they can be constructed as in Germany, do not as yet out-perform "domestic" definitions in terms of their link with final variables. However, CBHs are experiencing a rapid growth which suggests that they may become increasingly important for monetary analysis in the future.

As regards the definition of the scale variable in the demand equation, real GNP or GDP are selected depending on the availability. Wealth would be a natural candidate for inclusion in the specification, as portfolio considerations, in addition to transaction purposes, underlie the demand for broad money. Unfortunately, reliable data on both financial and total wealth are not available for all countries and thus neither variable can be computed at the area level. The price level variable is the GNP/GDP deflator corresponding to the scale variable.

The three-month interbank rate (R_I) and the secondary-market return on government bonds (R_B) are selected as short and long-term

---

7 The most comprehensive data base (BIS international banking statistics) provides quarterly series, which are not fully consistent before 1983 and cannot be added straightforwardly to national money stocks in each EC country because of double-counting problems.

rates, respectively, to proxy the return on alternative domestic financial assets. The own return on several national broad aggregates is not available and therefore could not be computed at the EC level. In many countries (e.g. France), a large part of the assets included in broad aggregates are non-interest bearing or are remunerated either at a fixed or a sluggish interest rate. Hence, the estimated impact on broad money demand of changes in the returns on financial assets can be expected not to be significantly different, whether the own return is included or not in the equation. However, as deregulation and innovation proceed, an increasing share of broad aggregates should be remunerated at floating rates and efforts to overcome this measurement problem constitute an obvious area for further work. In any case, R_I should not be expected to proxy the own return in the 1980s but rather the whole range of alternative assets, given its pivotal role for the entire term structure in many EC countries. Finally, foreign (third-currency) variables may be proxied by the US interest rates (defined consistently with R_I and R_B) and by the dollar/ECU exchange rate.

2. Aggregation at ERM-level

The method to convert quantity variables (money and income) into a common currency is controversial, as shown by the fact that in the literature different choices are made. Three alternative rates of conversion may be used: (i) current exchange rates, as applied for instance by Gray et al. (1976); (ii) fixed base-period exchange rates, as preferred by McKinnon (1982), Spinelli (1983) and Bekx and Tullio (1987); (iii) Purchasing Power Parity (PPP) rates selected by Kremers and Lane (1990).[8]

In this paper, the emphasis is placed on the first method. This corresponds to the prior that money balances are demanded, both for transaction and store-of-value purposes, with reference to their current and future actual purchasing power, which, in terms of foreign goods and assets, is appropriately measured by market exchange rates. Furthermore, current exchange rates provide a consistent market valuation across countries of the stocks of financial assets. Finally,

8 The methods turn out to be similar if nominal exchange rates are stable ((i) and (ii)); inflation rates are equal ((ii) and (iii)); and PPP holds continuously ((i) and (iii)).

intra-EC spillover effects are transmitted through variables expressed at current exchange rates.

However, in order to check the robustness of the conclusions obtained, the empirical analysis was repeated using money and income variables defined by the second method, with 1987 (last general realignment) as the base year. The third method is not used as it implies that the measurement of quantity variables, and hence the specification of money demand, is invariant to changes in relative prices of domestic and foreign goods. This feature seems to contradict the ultimate motive to hold money: to buy (domestic and foreign) goods, either in the near future (transactions) or at a later, possibly uncertain, date (precautionary and portfolio motivations). However, it may be worth noting that over the sample departures of current exchange rates from PPP rates are relatively small for all countries, except for the United Kingdom and, to a lesser extent, for Spain.

In the case of variables expressed in the same unit, such as indexes and rates, aggregation requires weighted averaging of national data. Prices (inflation) are weighed by current GNP/GDP (consistently converted at current or fixed-base exchange rates), and interest rates by the relative shares of EC currencies in the ECU basket after the 1989 revision. Following the previous EC-wide studies, the latter weights are used for financial variables, on the grounds that they can be expected to reflect more appropriately the role of national currencies and financial markets.

The main variables are shown in Charts 1 to 3.

IV. EMPIRICAL RESULTS

The presentation of empirical results, which follows the steps outlined in Section II, focuses on the money demand expressed in terms of variables converted into ECU at current exchange rates. Findings obtained for the variables converted at 1987 rates are reported in Appendix 2. The longest sample available spans from the first quarter of 1977 to the end of the third quarter of 1990. As an additional check of the robustness of the results, the analysis is replicated for the period starting with the EMS up to the end of 1989.

CHART 1: ERM REAL MONEY
INDEX (87Q1 = 100)

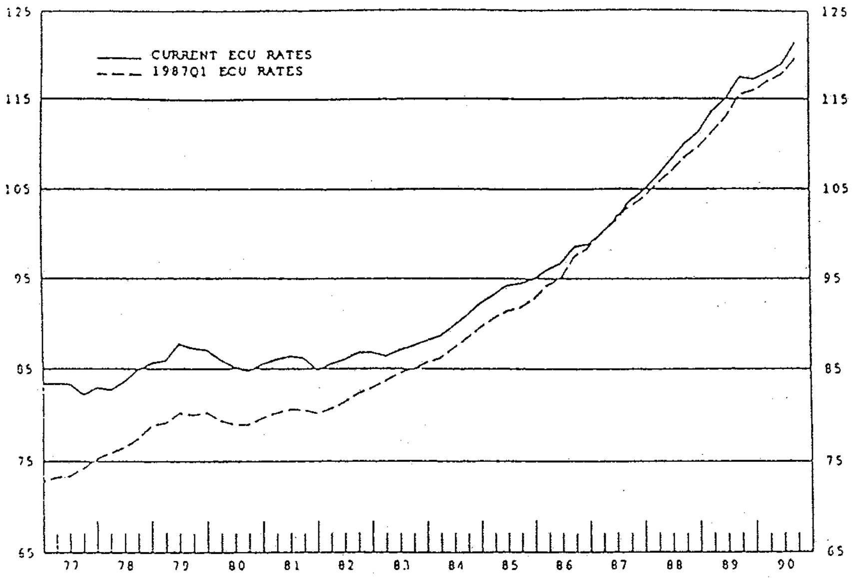


CHART 2: ERM VELOCITY
INDEX (87Q1 = 100)

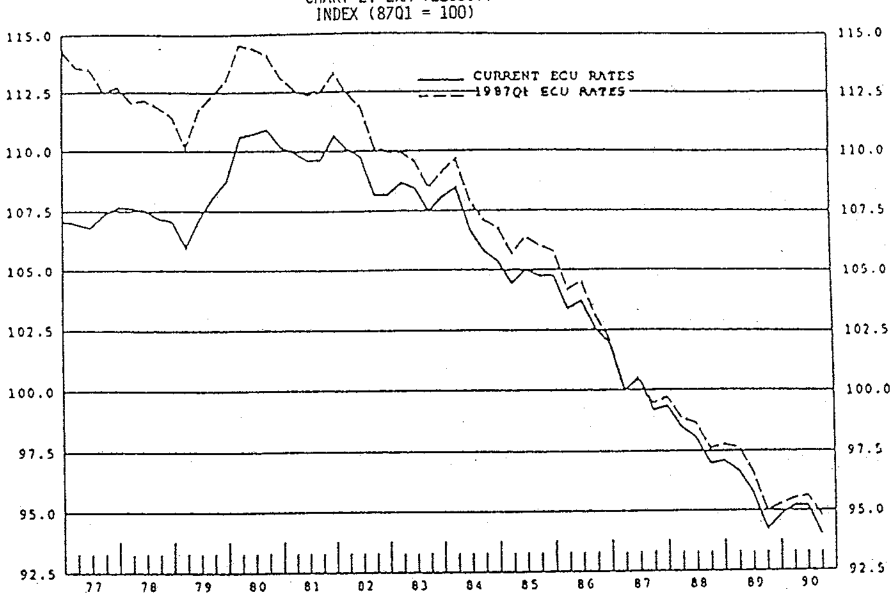
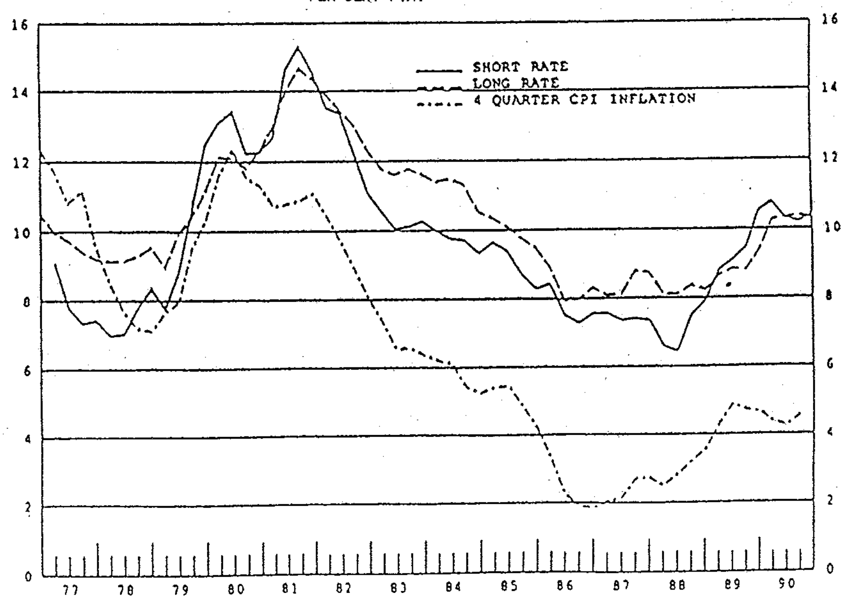


CHART 3: ERM INTEREST RATES AND INFLATION
PER CENT P.A.



1. Stationarity tests

As is often the case, the evidence for stationarity is not clear-cut, when tests are performed for different sub-samples and allowing or not for the presence of a constant and a trend.[9] Nevertheless, the overall findings tend to support the hypothesis that all relevant ERM-wide variables are I(1), i.e. integrated of order one, as is generally the case at the country level.

However, when applied to the first difference of the log of real income (y), the Phillips-Perron tests for both samples and conversion methods do not reject the hypothesis of the presence of a constant. The same applies to the first difference of the log of real money (m/p) for the presence of a constant and a trend. These results suggest that both variables could be I(1) around, respectively, a linear and non-linear trend. However, as hinted by Chart 1 and by the results of stationarity tests on sub-samples, real money could be I(1) around a trend starting only in the mid-1980s and not before. These findings suggest testing for the relevance of a segmented trend in the co-integrating vector.[10]

2. Co-integration and long-run estimates

A simple trivariate model with real money and real income (both converted at current ECU rates or at 1987 rates) and the interbank interest rate co-integrates, as shown by ADF test statistics and broadly confirmed by the Johansen procedure (see top of Table 2 and Appendix 2, Table A). This promising result, nevertheless, requires further testing.

Firstly, in line with the results of the stationarity tests for income and money, the relevance of a trend throughout the whole sample has to be assessed. Adding a linear trend spanning the entire

---

9 The results of stationarity tests are available from the authors upon request.

10 The resort to (non-linear) trends is not uncommon in national money demand studies for EC countries; see, for example, Muscatelli and Papi (1990) and Dolado and Escrivà (1991). On the other hand, preliminary analysis of detrended series was not promising, as detrending removed "too much" variability. For a discussion on the role of segmented trends in co-integration analysis see, e.g., Rappoport and Reichlin (1989) and Perron (1990).

TABLE 2: COINTEGRATION ON EC-WIDE MONEY WITH CURRENT EXCHANGE RATE (OLS & JOHANSEN)

EQU. Nber	PERIOD (Number of observations)	EXPLANATORY VARIABLES					DW	DF	ADF	JOHANSEN PROCEDURE			
		y	R_I	Trend (seg.)	Δ_4 ecu	cst.		cf Tables DF, ADF and Phillips-Ouliaris,1990		Ho	Trace test	Lambda max test	▾ Test para. equal to OLS
1a	77.1/90.3 (55)	1.63	-.73	-	-	-7.4	.62	-3.7	-3.57 o	r=0 r<1 r<2	34.46 * 18.44 * 5.37	16.01 13.07 5.37	6.11*
1b	79.1/89.4 (44)	1.74	-.50	-	-	-8.9	1.15	-5.5 **	-4.13 *	r=0 r<1 r<2	34.80 * 17.19 o 4.40	17.61 12.78 4.40	4.85
1c	77.1/90.3 (55)	1.28	-.65	.45	-	-2.8	1.14	-4.94 **	-4.92 **	r=0 r<1 r<2	38.86 ** 18.05 * 4.54	20.82 o 13.51 o 4.54	.82
1d	79.1/89.4 (44)	1.33	-.61	.44	-	-3.4	1.26	-4.69 *	-4.31 *	r=0 r<1 r<2	38.57 ** 18.42 * 4.35	20.15 o 14.07 o 4.35	3.37
1e	77.1/90.3 (55)	1.29	-.72	.46	.018	-2.9	1.22	-5.16 *	-4.84 DF ** P.O. *	r=0 r<1 r<2 r<3	59.65 ** 33.15 * 14.43 1.54	26.50 o 18.72 12.88 1.54	.91
1f	79.1/89.4 (44)	1.33	-.66	.43	.012	-3.4	1.31	-4.83 *	-4.29 DF * P.O. o	r=0 r<1 r<2 r<3	80.87 ** 38.07 ** 13.79 5.80	42.80 ** 24.29 * 7.98 5.80	5.67

o / * / ** : Test statistic above 10% / 5% / 1% critical value

▾ Value of the likelihood ratio test statistic that estimates from "Johansen" are equal to OLS estimates of parameters.

Example of first-stage OLS equation : $m\text{-}p = cst + \gamma\, y + \beta_1 R_I + \beta_2 \Delta_4 ecu + \beta_3$ Segm. Trend sample to the co-integrating vector has virtually no effect. In particular it does not modify the income elasticity, which appears to be high: about 1.6, for all samples and methods. Although there exist no strong priors for assuming a unitary income elasticity, especially in the case of a broad aggregate, estimates from national equations generally yield lower values. By contrast, the introduction of a (segmented) trend starting from the second half of the 1980s - as hinted by the results of the stationarity tests - reduces the income elasticity to about 1.3, while co-integration is confirmed. This estimate is not inconsistent with the values exceeding unity obtained in national studies for Italy (Muscatelli and Papi, 1990), Germany (Schmid and Herrman, 1991) and France (de Bandt, 1991) and with the prior that broad money is a superior good. Moreover, there are economic reasons suggesting that this trend may be capturing the effects of other determinants of money demand, missing in the trivariate model. In the second half of the 1980s, the acceleration in financial innovation and deregulation enhanced the liquidity and increased the return on assets included in broad aggregates, raising their share in the portfolio allocation of wealth (which cannot be computed at the area level) and leading to a steady decline in velocity over that period, shown in Chart 2.

The Johansen method confirms the co-integration result and offers support to the hypothesis that the co-integrating vector is unique,[11] owing both to the results of the trace tests and to the failure to reject the restriction that the co-integrating vector is equal to the one estimated by first-stage OLS. The robustness of these results is also corroborated by their stability across samples and across estimation techniques. In particular, equality restrictions with the long-run elasticities derived from the single-stage ECM are not rejected, as shown in Table 3.

Secondly, also the interest rate semi-elasticity results are very stable across samples and specifications (see Table 2), estimation techniques (see Table 3) and conversion methods (see Appendix 2). The estimates lie rather on the low side (around -0.7) of

---

11 Test statistics were compared with the critical values tabulated in Johansen and Juselius (1990). These values, however, do not take into account the presence of a segmented trend in the co-integrating vector.

TABLE 3: TESTS ON LONG-RUN ESTIMATES (CURRENT EXCHANGE RATES) (1)

EQUATION NUMBER (SAMPLE)	PROCEDURES	TEST VALUES (CORRESPONDING LONG-RUN ESTIMATES) y		R		Δ_4 ecu		JOINT TESTS
1c	Johansen		(1.28)		(-0.75)			
		} 0.00		} 0.81		-		} 0.82
	First stage		(1.28)		(-0.65)			
(long)		} 2.18		} 0.50		-		} 1.07
	Single stage		(1.39)		(-0.72)			
1d	Johansen		(1.41)		(-0.13)			
		} 0.18		} 3.36		-		} 3.37
	First stage		(1.33)		(-0.61)			
(short)		} 0.55		} 0.53		-		} 0.86
	Single stage		(1.45)		(-0.75)			
1e	Johansen		(1.28)		(-0.61)		(0.008)	
		} 0.00		} 0.41		} 0.58		} 0.91
	First stage		(1.29)		(-0.72)		(0.018)	
(long)		} 1.47		} 0.20		} 0.00		} 0.47
	Single stage		(1.37)		(-0.77)		(0.018)	
1f	Johansen		(1.28)		(-0.92)		(0.010)	
		} 0.24		} 1.83		} 0.00		} 5.67
	First stage		(1.33)		(-0.66)		(0.012)	
(short)		} 0.04		} 0.68		} 0.65		} 0.57
	Single stage		(1.36)		(-0.87)		(0.041)	

(1) Tests on the restriction that Johansen or single-stage estimates are respectively equal to first-stage estimates.
For Johansen, use of likelihood ratio test on estimates obtained from the eigenvector corresponding to the highest eigenvalue, based on the assumption of a unique co-integrating vector.
For single stage, use of F-test.

the range of semi-elasticities reported in national studies for European countries, except for the United Kingdom, where Hall et al. (1990) could not find any significant response of money demand to interest rates. The results obtained for the three-month interbank rate (R_I) are not significantly modified when the bond rate (R_B) is alternatively used, while their simultaneous insertion raises severe multicollinearity problems (see the plot of the two variables in Chart 3). However, R_I is preferred to R_B and maintained throughout the analysis on the grounds of its superior explanatory power in the dynamic equations.

Thirdly, all attempts to introduce supplementary variables, such as inflation or the return on foreign assets (US interest rates) in the co-integrating vector turn out to be disappointing, except for the four-quarter rate of change in the dollar/ECU exchange rate. The relevance of this variable is confirmed for different samples, techniques and conversion methods, although not overwhelmingly (cf. bottom of Table 2, Table 3 and Appendix 2). Its positive impact shows that an appreciation of the ECU increases the demand for ERM money, as it makes ERM assets more attractive than alternative assets denominated in third currencies (US dollars). The (imperfect) substitutability between monetary assets denominated in ERM and third currencies is shown to play a role in the determination of the ERM-wide money demand. However, the elasticity to the return on assets denominated in US dollars appears to be lower than that to the return on alternative assets of the ERM area.

3. ECM and short-run dynamics

Table 4 (as well as Table C in Appendix 2) reports the results on the ECM (single-stage or second-stage), which have been estimated on the basis of the co-integrating vectors, including the segmented trend and both including and excluding the dollar/ECU exchange rate. The application of OLS to ECMs is only efficient if the explanatory variables are weakly exogenous (see Engle et al., 1983). As weak exogeneity implies that the co-integrating vector can only appear in the equation determining money, suitable restrictions on the matrix of "α-weights" obtained from the Johansen procedure were tested. The evidence did not reject weak exogeneity for the long sample with variables converted at current exchange rates and without

TABLE 4 : ERROR-CORRECTION MODEL ON EC-WIDE MONEY WITH CURRENT EXCHANGE RATE (second stage or single stage)

STAGE	Eq. Number cf Tab. 2 (sample)	LAGGED LEVELS (& error-corr. term)					DIFFERENTIATED VARIABLES [lag]				TREND (seg.)	R^2	S.E x 100	DW	TESTS NOT PASSED (1)
		$\hat{\varepsilon}_{-1}$	m_{-1}	y_{-1}	R_{-1}	$\Delta_4 ecu_{-1}$	$\Delta m[-i]$	$\Delta y[0]$	$\Delta R[-1]$	$\Delta\Delta_4 p[0]$					
SINGLE	2c (long)		-.423 (4.17)	.589 (4.45)	-.306 (4.32)	-	.319 [2] (2.89)	.624 (6.1)	-.26 (2.2)	-.26 (1.9)	.16	.74	.49	1.51	LM(1-10)*
	2d (short)		-.292 (2.49)	.424 (2.45)	-.218 (2.78)	-	.324 [2] (2.31)	.522 (3.96)	-.45 (3.81)	-	.116	.71	.51	1.64	-
	2e (long)		-.444 (4.28)	.612 (4.53)	-.342 (4.27)	.008 (0.9)	.288 [2] (2.48)	.624 (6.1)	-.24 (1.99)	-.26 (1.86)	.16	.75	.49	1.55	LM(1-10)*
	2f (short)		-.312 (2.56)	.425 (2.44)	-.273 (2.72)	.0127 (1.11)	.272 [2] (1.76)	.497 (3.62)	-.39 (3.12)	-	.10	.72	.52	1.82	-
SECOND	2'c (long)	-.463 (4.76)					.234 [1] (2.69) .413 [2] (4.52)	.643 (6.98)	-.27 (2.4)	-.29 (2.12)		.73	.49	1.68	-
	2'd (short)	-.372 (3.35)					.466 [2] (4.69)	.602 (5.3)	-.49 (4.45)	-		.68	.52	1.5	-
	2'e (long)	-.443 (4.41)					.217 [1] (2.43) .385 [2] (4.16)	.64 (6.77)	-.26 (2.31)	-.27 (1.91)		.72	.50	1.63	-
	2'f (short)	-.371 (3.31)					.45 [2] (4.53)	.60 (5.28)	-.47 (4.3)	-		.68	.52	1.5	-

(1) Diagnostic tests performed: see Table 1   * / ** test statistic above 5% / 1% critical value

Example of single-stage equation : $\Delta(m-p) = cst. + a_o m_{-1} + a_1 y_{-1} + a_2 R_{-1} + a_3 \Delta_4 ecu_{-1} + b\Delta m_{-2} + c\Delta y + d\Delta R_{-1} + e\Delta\Delta_4 p + f$ Seg. Trend

Example of second-stage equation : $\Delta(m-p) = cst. + a\,\hat{\varepsilon}_{-1} + b\Delta m_{-2} + c\Delta y + d\Delta R_{-1} + e\Delta\Delta_4 p$

the ECU as an explanatory variable, while the results were less clear-cut in other cases.[12]

All equations track well the quarterly rate of growth of the EC-wide broad aggregate. The standard error of the estimate is about or below 0.5 percentage point, comparing favourably with the best single-country studies and with Kremers and Lane's result (1990) for a narrow aggregate, 0.8%. The quality of the statistical performance of the equations is confirmed by the thorough battery of diagnostic tests (listed in Table 1) which are passed, with the only exception of the Lagrange multiplier test of order 1 to 10 for the single-stage ECM on the long sample. Parameters appear to be stable over time, as revealed by out-of-sample Chow tests computed for breaks at the end of 1986, 1987 and 1988 and by the results of recursive estimation, omitted here for the sake of brevity. An illustration of the dynamic properties is given by Charts 4 and 5, which report examples of recursive estimates and out-of-sample simulations.

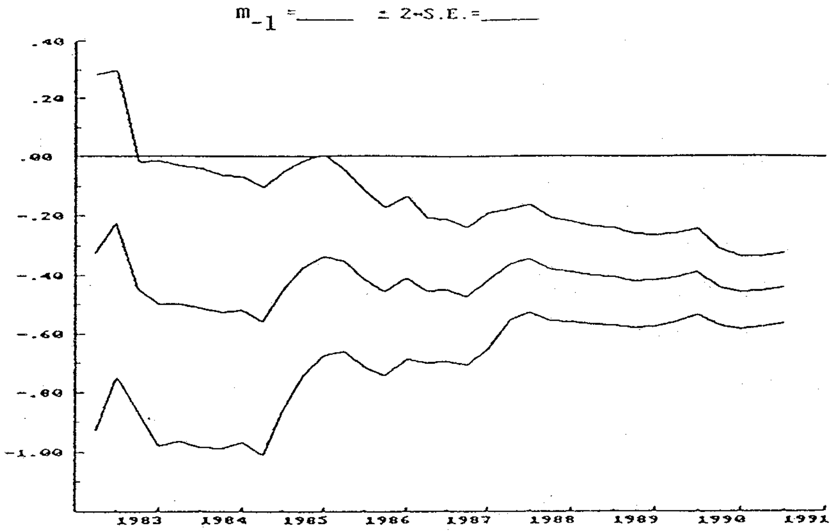
The pattern of responses that emerges from the results is very similar across specifications and samples for the variables converted at current exchange rates. In cases of departures from equilibrium, about 30% to more than 40% of the shock is corrected within one quarter.[13] As far as the dynamics of the desired real money balances are concerned, it depends positively on current changes in real income (with a coefficient of about 0.6) and negatively on the lagged change in interest rate, with a coefficient around -0.25 for the longest sample, where the contemporaneous change in the rate of CPI inflation also has an effect of the same magnitude. As shown in Appendix 2, the dynamics of the ECM for variables converted at the 1987 exchange rate are broadly similar, although the speed of adjustment is slightly lower and perhaps less consistent across specifications and samples.

---

12 The results of these tests, as well as further details on the evidence from the Johansen procedure, are available from the authors upon request.

13 This finding is in line with the indication about the average speed of adjustment which is supplied by the estimates of the "α-weights" obtained from the Johansen procedure, which range from -0.31 to -0.53.

Chart 4 : RECURSIVE ESTIMATES OF (LAGGED) MONEY AND INCOME COEFFICIENTS
(single-stage equation 2 - long sample)



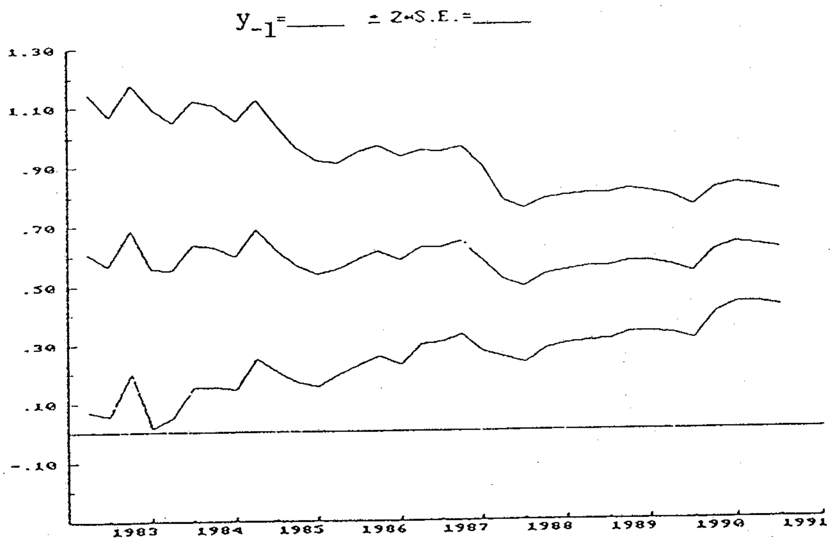
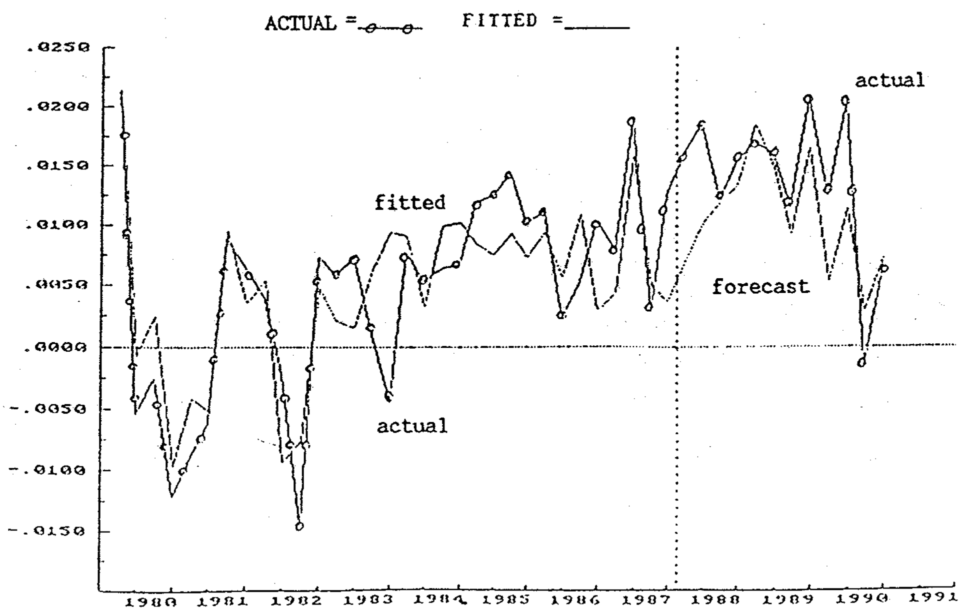


Chart 5: RATE OF GROWTH OF REAL BROAD ERM MONEY: ECM ESTIMATES
(second-stage equation 2´e - Chow test (12,27) : 0.82)



Finally, it is worth comparing the characteristics of the broad money demand equation for the ERM area with the results obtained for the U.S. following the co-integration/ECM approach. Most recent studies (e.g. Boughton, 1991; Boughton and Tavlas, 1991; Miller, 1991; Hafer and Jansen, 1991) find co-integration between M2 (results for M1 appear less robust), real income and a short-term interest rate, with a long-run coefficient for the latter variable which is very close to the value obtained for the ERM area. The dynamic performance of most equations is very similar to the findings for the ERM area in terms of accuracy (standard errors are around 0.5%), while the speed of adjustment to shocks is lower, with less than 20% of the shock corrected in the first quarter.

V. CONCLUSIONS

The analysis presented in this paper indicates that, over the 1980's, the aggregate demand for broad money at the ERM level is a stable function of income and interest rate (and the rate of change in the ECU-dollar exchange rate). The robustness of this result is supported by its minimal sensitivity to the choice of alternative samples, estimation techniques and methods of converting aggregate variables into ECU. Furthermore, the statistical performance of the dynamic equations compares favourably with both national studies and previous work on the area-wide demand for narrow money.

The long-run estimates show that the ERM-wide broad money demand has a relatively low elasticity to interest rates (around -0.7) and an income elasticity around 1.3. The latter feature together with the (segmented) trend which appears in the co-integrating vector suggest that, for the ERM area as a whole, monetary expansion at a rate higher than that of nominal income growth is not incompatible with the maintenance of nominal stability.

If corroborated further, the econometric evidence summarised above has two important implications. First, the degree of economic and financial integration within the ERM countries is already sufficient for an ERM-aggregate money demand to be a valuable tool - at least as powerful as single-country equations at the national level - for analysing and interpreting monetary developments within the Community. Secondly, in view of the role played by broad aggregates in the formulation of monetary policy in several European countries, the existence of a stable and well-behaved money demand function at the ERM level provides the necessary technical condition for a broad monetary aggregate to assist in the conduct of an ERM-wide monetary policy based on Community variables.

Although "Goodhart's law" hovers darkly and hence no econometric evidence can ensure that money demand stability will persist when exploited for monetary control, further examination of the results of this paper is clearly desirable to enhance the reliability of the above conclusions as European integration progresses. In particular, efforts towards the calculation of an adequate proxy for financial and real wealth and towards the definition of the own-return on an ERM-wide broad monetary aggregate stand as promising areas for future research together with the analysis of the role of monetary cross-border holdings.

REFERENCES

BANERJEE A., J. DOLADO, D. HENDRY and G. SMITH, (1986), "Exploring Equilibrium Relationships in Econometrics Through Static Models: Some Monte-Carlo-Evidence", Oxford Bulletin of Economics and Statistics, Vol. 48, pp. 253-277.

BEKX P. and G. TULLIO, (1989), "A Note on the European Monetary System and the Determination of the DM-Dollar Exchange Rate", Cahiers Economiques de Bruxelles, No. 123, pp. 329-343.

BOUGHTON J.M., (1991a), "Long-Run Money Demand in Large Industrial Countries", IMF Staff Papers, Vol. 38, pp. 1-32.

BOUGHTON J.M., (1991b), "Money Demand in Five Major Countries: Estimating and Interpreting Error Correction Models", in M.P. Taylor (ed.), Money and Financial Markets, Basil Blackwell, London

BOUGHTON J.M. and G.S. TAVLAS, (1991), "What Have We Learned About Estimating the Demand for Money?: A Multi-Country Evaluation of Some New Approaches", IMF Working Paper, No. 91/16, February.

DE BANDT O., (1991) "Politique monétaire et agregats de monnaie en France", Bulletin Trimestriel de la Banque de France, 78, pp. 79-93.

DICKEY D. and W. FULLER, (1981), "Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root", Econometrica, Vol. 49, pp. 1057-1072.

DOLADO J.J. and J.L. ESCRIVA, (1990) "The Demand for Money in Spain: Broad Definitions of Liquidity", Bank of Spain, Documento de Trabajo no. 9107.

DOLADO J.J., T. JENKINSON and S. SOSVILLA-RIVERO, (1990), "Cointegration and Unit Roots", Journal of Econometric Surveys, Vol. 4, pp. 249-273.

ENGLE R.F. and C.W.J. GRANGER, (1987), "Cointegration and Error Correction: Representation, Estimation and Testing", Econometrica, Vol. 55, pp. 251-276.

ENGLE R.F. and B.S. YOO, (1987), "Forecasting and Testing in Cointegrated Systems", Journal of Econometrics, Vol. 35, pp. 143-159.

FASE M.M.G. and C.C.A. WINDER, (1990), "The Demand for Money in the Netherlands Revisited", De Economist, n. 138, pp. 276-301.

GOODHART C. (1989), "International Financial Interlinkages", LSE Financial Markets Group Special Paper Series, n. 21

GRANGER C.W.J. (1983), "Cointegrated Variables and Error-Correcting Models", University of California Working Paper 83-13, San Diego

GRANGER C.W.J. and A. WEISS (1983), "Time Series Analysis of Error Correction Models", in S. Karlin, T. Amemiya and L. Goodman (eds.), Studies in Economic Time Series and Multivariate Statistics, Academic Press, New York

GRAY M.R., R. WARD and G. ZIS, (1976), "The World Demand for Money Function: Some Preliminary Results", in M. Parkin and G. Zis (eds.), Inflation in the World Economy, Manchester University Press.

HALL S.G., S.G.B. HENRY and J.B. WILCOX, (1990), "The Long-Run Determination of UK Monetary Aggregates" in S.G.B. Henry and K.D. Patterson (eds.), Economic Modelling at the Bank of England, Chapman and Hall, London.

HENDRY D.F., (1980), "Predictive Failure and Econometric Modelling in Macroeconomics: The Transaction Demand for Money", in OMEROD P. (ed.), Economic Modelling, Heinemann Educational Books, London.

HENDRY D.F., (1985), "Monetary Economic Myth and Econometric Reality", Oxford Review of Economic Policy, Vol. 1, pp. 72-84.

HENDRY D.F., (1986), "Econometric Modelling with Cointegrated Variables: An Overview", Oxford Bulletin of Economics and Statistics, Vol. 48, pp. 201-212.

HENDRY D.F., (1989), "PC-GIVE: An Interactive Econometric Modelling System", Institute of Economics and Statistics, University of Oxford, Oxford.

HURLEY M. and GUIOMARD C., (1990), "The determinants of Money Demand in Ireland 1972-1989: Rounding up the Usual Suspects", Economic and Social Review.

HYLLEBERG S., R.F. ENGLE, C.W.J. GRANGER and B.S. YOO, (1990), "Seasonal Integration and Cointegration", Journal of Econometrics, Vol. 44. pp. 215-238.

JOHANSEN S., (1988), "Statistical Analysis of Cointegrating Vectors", Journal of Economic Dynamics and Control, Vol. 12, pp. 231-254.

JOHANSEN S. and K. JUSELIUS, (1990), "Maximum-Likelihood Estimation and Inference on Cointegration - With Applications to the Demand for Money", Oxford Bulletin of Economics and Statistics, Vol. 52, pp. 169-210.

KREMERS J.M. and T.D. LANE, (1990), "Economic and Monetary Integration and the Aggregate Demand for Money in the EMS", IMF Staff Papers, Vol. 37, pp. 777-805.

MCKINNON R.I., (1982), "Currency Substitution and Instability in the World Dollar Standard", American Economic Review, Vol. 72, pp. 320-333.

MILLER S., (1991), "Money Dynamics: An Application of Cointegration and Error-Correction Modelling", Journal of Money, Credit and Banking, Vol. 23, pp. 139-154.

MUSCATELLI V.A. and L. PAPI, (1990), "Cointegration, Financial Innovation and Modelling the Demand for Money in Italy", Manchester School, Vol. 58, pp. 242-259.

OSBORN D., A. CHUI, J. SMITH and C. BIRCHENHALL, (1988), "Seasonality and the Order of Integration for Consumption", Oxford Bulletin of Economics and Statistics, Vol. 50, pp. 361-377.

PERRON P., (1990), "Further Evidence for Breaking Trend Functions in Macroeconomic Variables", Princeton University Working Paper No. 350.

PHILLIPS P. and P. PERRON, (1988), "Testing for a Unit Root in Time Series", Biometrika, Vol. 75, pp. 335-346.

PHILLIPS P. and S. OULIARIS, (1990): "Asymptotic Properties of Residual Based Tests for Cointegration", Econometrica, Vol. 58, pp. 165-193.

RAPPOPORT P., and L. REICHLIN, (1989), "Segmented Trends and Non-Stationary Time Series", The Economic Journal, Supplement, Vol. 99, pp. 168-177.

SCHMID P. and HERRMAN H. (1991), "Federal Republic of Germany: Old and new problems posed by a target-guided monetary policy", in "The Orientation of Monetary Policy and the Monetary Policy Decision-Making Process", Bank for International Settlements, April, pp. 68-86.

SPINELLI, F., (1983), "Currency Substitution, Flexible Exchange Rates, and the Case for International Monetary Co-Operation: Discussion of a Recent Proposal", IMF Staff Papers, Vol. 30, pp. 755-783.

STOCK J., (1987), "Asymptotic Properties of Least Square Estimators of Cointegrating Vectors", Econometrica, Vol. 55, pp. 1035-1056.

Appendix 1

DATA SOURCES

1. Broad monetary aggregates (1991 definitions)

All aggregates, obtained from national central banks, are quarterly averages of monthly, seasonally-adjusted figures, except for UK M4 which is quarterly, seasonally-adjusted. Belgium: M2H; Denmark: M2 less Treasury Bills; Germany: M3 old Federal Republic; Spain: ALP plus foreign-currency deposits less Treasury Bills; France: M3 (new definition) less 5 years' non-negotiable certificates; Ireland: M3; Italy: M2 less Buoni Postali Fruttiferi, plus Assegni Cambiari; Netherlands: M3; United Kingdom: M4 plus foreign-currency deposits.

2. Real income

Real income is measured by real GDP (BIS Data bank) for all countries except Denmark, Spain and the Netherlands (national central bank sources) and for Belgium and Ireland (OECD Economic outlook: yearly data interpolated), where GNP is used.

3. GDP/GNP implicit price deflators

All deflators derive from the same source as the corresponding seasonally-adjusted income variables.

4. Short-term interest rates

All short-term interest rates are quarterly average of monthly average figures (three-month interbank, BIS Data bank), except for Ireland, where quarterly averages of end-month values were provided by the Central Bank of Ireland.

5. Long-term interest rates

For long-term rates, BIS data for secondary markets of government or public sector bonds are used, except for Spain and Ireland (central bank sources). Rates are quarterly average of monthly average figures for Germany, Spain, Italy and the Netherlands, whereas quarterly averages of end-month values are used for Belgium, Denmark, France, Ireland and the United Kingdom.

6. ECU exchange rates

All ECU exchange rates are quarterly averages of theoretical spot rates. For the period before 1979, artificial rates were constructed using the 1979 set of weights in the basket.

TABLE A : COINTEGRATION ON EC-WIDE MONEY WITH 1987 EXCHANGE RATE (OLS & JOHANSEN) APPENDIX 2

EQU. Nber	PERIOD (Number of observations)	EXPLANATORY VARIABLES y	R I	Trend (seg.)	Δ ecu 4	cst.	DW	DF (cf Tables DF, ADF and Phillips-Ouliaris,1990)	ADF	JOHANSEN PROCEDURE Ho	Trace test	Lambda max test	▾ Test para. equal to OLS
1a	77.1/90.3 (55)	1.66	-.62	-	-	-7.8	.82	-4.12 **	-4.05 ADF ** P.O. *	r=0 r<1 r<2	62.62 ** 20.08 * 1.80	42.54 ** 18.28 * 1.80	30.46**
1b	79.1/89.4 (44)	1.74	-.41	-	-	-8.9	1.17	-5.08 **	-4.65 **	r=0 r<1 r<2	51.12** 15.42 2.92	35.70** 12.50 2.92	20.15**
1c	77.1/90.3 (55)	1.50	-.56	.28	-	-5.6	1.18	-4.81 *	-4.48 *	r=0 r<1 r<2	73.22 ** 25.44 * 10.51 *	47.77 ** 14.94 * 10.51 *	16.28**
1d	79.1/89.4 (44)	1.53	-.54	.26	-	-6.1	1.21	-4.48 *	-4.76 **	r=0 r<1 r<2	39.34 ** 17.87 * 7.17	21.46 * 10.71 7.17	6.8*
1e	77.1/90.3 (55)	1.50	-.70	.28	.028	-5.6	1.35	-5.31 *	-4.89 **	r=0 r<1 r<2 r<3	111.55 ** 56.37 ** 26.20 ** 9.08 *	58.18 ** 27.17 ** 17.12 * 9.08 *	25.26**
1f	79.1/89.4 (44)	1.53	-.65	.26	.023	-6.0	1.33	-4.76 *	-4.66 *	r=0 r<1 r<2 r<3	90.61 ** 52.70 ** 23.45 ** 2.56	37.91 ** 29.25 ** 20.88 ** 2.56	7.21

* / ** : Test statistic above above 5% / 1% critical value

▾ Value of the likelihood ratio test statistic that estimates from "Johansen" are equal to OLS estimates of parameters

TABLE B: TESTS ON LONG-RUN ESTIMATES (FIXED EXCHANGE RATES) (1) APPENDIX 2

EQUATION NUMBER (SAMPLE)	PROCEDURES	TEST VALUES (CORRESPONDING LONG-RUN ESTIMATES)			JOINT TESTS
		y	R	Δ_4 ecu	
1c	Johansen	(1.45)	(-0.24)		
		} 0.08	} 16.13**	-	} 16.28**
	First stage	(1.50)	(-0.56)		
(long)		} 1.06	} 3.83	-	} 2.47
	Single stage	(1.60)	(-0.96)		
1d	Johansen	(1.26)	(-0.26)		
		} 1.80	} 4.70*	-	} 6.80*
	First stage	(1.53)	(-0.54)		
(short)		} 0.00	} 3.76	-	} 2.45
	Single stage	(1.53)	(-1.19)		
1e	Johansen	(1.34)	(-0.31)	(0.001)	
		} 0.86	} 19.96**	} 3.56	} 25.26**
	First stage	(1.50)	(-0.70)	(0.028)	
(long)		} 0.92	} 7.09*	} 1.89	} 2.46
	Single stage	(1.35)	(-1.40)	(0.100)	
1f	Johansen	(1.12)	(-0.86)	(0.027)	
		} 2.96	} 1.21	} 0.08	} 7.21
	First stage	(1.53)	(-0.65)	(0.023)	
(short)		} 3.11	} 4.05	} 3.73	} 1.61
	Single stage	(1.21)	(-1.24)	(0.110)	

(1) Tests on the restriction that Johansen or single-stage estimates are respectively equal to first-stage estimates.
For Johansen, use of likelihood ratio test on estimates obtained from the eigenvector corresponding to the highest eigenvalue, based on the assumption of a unique co-integrating vector.
For single stage, use of F-test.

* Test statistic above 5% critical value.
** Test statistic above 1% critical value.

APPENDIX 2

TABLE C : ERROR-CORRECTION MODEL ON EC-WIDE MONEY WITH FIXED EXCHANGE RATE (second stage or single stage)

STAGE	Eq. Number cf Tab. A (sample)	LAGGED LEVELS (& error-corr. term)					DIFFERENTIATED VARIABLES [lag]					TREND (seg.)	R^2	S.E x 100	DW	TESTS NOT PASSED (1)
		$\hat{\varepsilon}_{-1}$	m_{-1}	y_{-1}	R_{-1}	$\Delta_4 ecu_{-1}$	Δ m[-i]	Δ y[-i]	Δ R[-1]	$\Delta\Delta_4$ ecu[-i]	$\Delta\Delta_4$ p[-i]					
SINGLE	2c (long)		-.170 (1.91)	.272 (2.00)	-.164 (3.46)	-	.333 [2] (2.53)	.209 [0] (1.88)	-.259 (2.56)	-	-.379 [0] (2.81)	.05	.69	.401	1.65	-
	2d (short)		-.129 (1.85)	.198 (1.86)	-.153 (2.91)	-	.427 [2] (3.21)	-	-.239 (2.26)	.018 [1] (1.80)	-	.06	.78	.355	1.84	-
	2e (long)		-.313 (1.65)	.177 (1.54)	-.183 (3.13)	.013 (1.74)	.281 [2] (2.17)	-	-.282 (2.95)	-	-.320 [0] (2.47)	.04	.69	.400	1.99	-
	2f (short)		-.190 (2.61)	.230 (2.18)	-.236 (3.25)	.021 (2.29)	.426 [2] (3.27)	-	-.231 (2.21)	-	.426 [0] (3.27)	.07	.80	.345	1.89	-
SECOND	2'c (long)	-.265 (3.39)					.266 [1] (2.33) .474 [2] (4.60)	.297 [0] (3.12)	-.315 (3.39)	-	-.311 [0] (2.38)	-	.66	.409	1.69	-
	2'd (short)	-.407 (4.19)					.370 [1] (3.20) .687 [2] (5.87)	.188 [0] (1.79) -.214 [1] (1.95) -.214 [2] (1.95)	-.296 (2.74)	-	-.520 [0] (3.50) -.286 [1] (1.66)	-	.77	.363	1.78	-
	2'e (long)	-.266 (3.28)					.218 [1] (2.25) .454 [2] (4.42)	.291 [0] (3.05)	-.310 (3.30)	-	-.307 [0] (2.34)	-	.65	.411	1.66	-
	2'f (short)	-.343 (4.60)					.260 [1] (2.89) .615 [2] (6.38)	.242 [0] (2.46)	-.273 (2.97)	.018 [1] (1.75) .019 [2] (2.06)	-.583 [0] (3.92)	-	.79	.350	1.79	-

(1) Diagnostic tests performed : see Table 1   * test statistic above 5% critical value   ** test statistic above 1% critical value