European Journal of Political Economy
Vol. 16 (2000) 587–609

European Journal of POLITICAL ECONOMY

Endogeneity in European money demand

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Received 1 March 1999; received in revised form 1 January 2000; accepted 1 June 2000

Abstract

European wide monetary aggregates constructed from pre-unification data cannot be used as evidence that money demand in the euro area is stable. To overcome the Lucas critique, we apply the standard foreign exchange rate model. Since the uncoordinated country specific money supply system is abolished, the increased comovement between local monetary aggregates leaves little room for a free ride on the law of large numbers. Current monetary policy decisions must be based on untested relations, and given 'the long and variable lags', we conclude that the road towards monetary stability is a non-activist steady money supply policy. © 2000 Elsevier Science B.V. All rights reserved.

JEL classification: E41; E52
Keywords: European money demand; Lucas critique

1. Introduction

The coming of the Euro has put new life into money demand research. A recent, but by no means exhaustive, survey by the European Monetary Institute (Browne et al., 1997) lists at least 15 articles that use artificially constructed monetary aggregates in the estimation of the European money demand function. This line of research started with the papers by Bekx and Tullio (1989) and

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PII: S0176-2680(00)00018-5

Kremers and Lane (1990). Other notable papers are Artis et al. (1993), Monticelli and Strauss-Kahn (1993), Cassard et al. (1994), Artis (1996), Monticelli (1996), Tullio et al. (1996), Wesche (1997), La Cour and MacDonald (1997), Spencer (1997), Fagan and Henry (1998) and Hayo (1999). From these studies, a remarkable consensus has emerged on the degree of stability of the European demand for money. Taking the residual standard error as a rough-and-ready indicator of stability, a typical European average money demand function beats the former German money demand, generally perceived as having been one of the world's most stable, by at least 30%. Furthermore, standard econometric stability tests on pre-Economic and Monetary Union (EMU) data fail to detect any sign of structural instability in European money demand functions.

On the basis of this apparent stability, these studies have been interpreted as providing support for the beneficial effects of monetary union. Monetary integration would stabilize the rather erratic monetary aggregates in Europe. This would make the life of European policy makers and monetary authorities a lot easier. Moreover, because of this stability, monetary targeting would be feasible. The Bundesbank, for example, has used this result as an argument in its campaign for the implementation of the old Bundesbank operational rules at the European level, rather than choosing for the pragmatic inflation targeting rule followed, for example, by the Bank of England.

In policy circles, the ostensible stability of European money demand has been met with cautious optimism, instead of a more critical too-good-to-be-true response. Doubts about whether these econometric exercises have been properly interpreted, however, are nagging. The area-wide average of individual money demand functions estimated for countries that did participate in the European Monetary System (EMS) may not be representative for the area-wide money demand in the EMU. The results based on the pre-EMU average are liable to the Lucas (1976) critique. Specifically, money demand equations are not exogenous to the institutional environment, since agents respond to the change in regime.

General concern that EMS-money demand may not be extrapolated to EMU has been voiced amongst others by Giovannini (1991). More concrete evidence for the non-applicability of the results has been given by Arnold (1997) on the basis of a comparison between regional money demands in the United States and national money demands in Europe. The residual cross-correlations between regions in the United States appear to be much higher than those between European nations. De Grauwe (1996) shows similar evidence for the regional differences in inflation between the German länder and the inflation differences in the EMS countries. This confirms the basic economic intuition that, inside a monetary union, monetary developments to a large extent run in parallel. For EMU, this implies that, since its start, the correlations between the former national money demands must have increased.

The increase in monetary comovement across the European nations and the resultant increased aggregate variability that we predict in this paper, is an application of the Hayekian argument that policy centralization enhances the variability of policy outcomes. The statistical analysis by Arnold (1997) is in this respect rather suggestive. In the current paper, we provide support for these statistical arguments by means of an economic analysis. Ex ante analysis of regime changes is notoriously difficult due to the Lucas critique. This is especially so if similar type of regime changes have been very rare. But we argue that the standard open economy model for exchange rate determination is well suited for handling the Lucas critique. In particular, the model is approximately structurally invariant under the regime switch from managed float to monetary union. By interchanging the endogenous and exogenous variables, the model can be used to predict the main features of future demand for money in Europe.

The foreign exchange rate model separates two effects. The first is that, during the pre-EMU years, idiosyncratic shocks in national money demands averaged out due to the law of large numbers. Hence, the variance of the average monetary aggregate will be lower than the variance of the constituent parts. Whether this averaging effect carries over to monetary union is doubtful. Some authors, e.g. Kremers and Lane (1990), have suggested that intra-European currency substitution may be behind the averaging effect. If currency substitution was an important phenomenon within Europe, monetary union would indeed increase money demand stability. But the empirical evidence shows intra-European currency substitution to be negligible (see Angeloni et al., 1992; Mizen and Pentecost, 1994; Bundesbank, 1995). The stability of European money demand thus has little to do with currency substitution, as argued more extensively in Arnold (1996). Absent currency substitution, the averaging effect will be lost due to unification. The pre-union variability of the average is an illusory predictor for the post-union variability of the area-wide monetary aggregate. All that can be said is that it provides a nice textbook example for what the Lucas critique is all about.

Instead of this illusory effect, we recognize the following endogenous response of European money demand to unification. As a result of the unitary monetary policy, we predict that the correlation in the movements of monetary aggregates across the union will increase dramatically. This second effect stems from the quantitative importance of the intra EU exchange rate variability vis-a-vis the variability of the other variables in the model. Because exchange rate variability is driven to zero by monetary unification, the exchange rate model implies that the covariance between local monetary aggregates increases by about the amount of the decline in exchange rate variability.

This endogenous response is akin to Goodhart's Law, which states that: '... any observed statistical regularity will tend to collapse once pressure is placed upon it for control purposes' (Goodhart, 1975, p. 5). Both deal with the effect of a regime-shift on money demand, but there is a major distinction. Goodhart's law concerns the conduct of monetary policy. Our effect pertains to how monetary union changes the comovement of monetary aggregates, irrespective of how actual policy is conducted.

The relevance of these two predictions, i.e. the loss of the law of large numbers effect and the increased correlation among monetary aggregates, for current monetary policy making by the ECB is as follows. The illusory findings of extreme European money demand stability wrongly suggest there is room for manoeuvre. These studies may have triggered the demands for an activist stance on monetary policy making during the first year of the ECB. But just as a regiment of soldiers that marches in lockstep across a bridge risks the collapse of the bridge, the higher correlations tend to destabilize European-wide money demand under activist monetary policy making. Thus, a higher degree of monetary harmonization in Europe does not automatically lead to extra monetary stability, as one might have wished. Instead, the amount of monetary prudence exercised by the ECB determines the amount of variability of the union wide monetary aggregates.

The monetary policy regime is partly shaped by the institutional design of the ECB. For example, the power balances within the board between local and EC appointees will be important for determining the policy decisions (see Von Hagen and Süppel, 1994). Of prime importance is the explicit goal of price stability. But there is of course room for independent policy making by the board of the ECB. The suggested extra room for manoeuvre may therefore lead to a more activist and hence variable monetary policy than is compatible with the stated objective of price stability. Unfortunately, due to the well-known ‘long and variable lags’ in the transmission of monetary policy decisions, it can take years before one can tell from the data that there is no such extra room for activism in the monetary union. Hence, the importance of arguments as provided in this paper to show that the pre-union data do not imply there is a benefit to monetary unification in that sense.

In Section 2, we identify the two effects discussed above. Section 3 considers the quantitative importance of the different effects. Section 4 concludes.

2. Analysis of monetary unification

The analysis is built on the Euler equations for international asset diversification and the local quantity equation. Consider the following two-period consumption-investment problem:

$$\max_{X, X_j, B, D} U(X) + \Sigma_j \pi_j W(X_j)$$

subject to

$$Y = PX + B + SD,$$

$$Y_j + RB + S_j ID = P_j X_j.$$

Here, X denotes consumption with price P, Y is nominal income, B are domestic bonds with gross returns R, D are foreign bonds with gross returns I denominated in foreign currency, and S is the foreign exchange rate. The states of the world are indicated by subscript j, and state probabilities are π_j. The expected utility function is time additive; $U(X)$ is the first period utility, and $\Sigma_j \pi_j W(X_j)$ is the second period expected utility.

Let $K_j = W_{xj}/U_x$ be the intertemporal marginal rate of substitution, or briefly the pricing kernel. The first-order conditions imply the following two equalities if there exists an interior solution:

$$E\left[K_j \frac{P}{P_j} R\right] = 1, \tag{1}$$

$$E\left[K_j \frac{P}{P_j} \frac{S_j}{S} I\right] = 1. \tag{2}$$

Denote foreign variables by a superscript star, except for foreign bonds and bond returns. Abroad a similar analysis yields:

$$E\left[K_j^* \frac{P^*}{P_j^*} I\right] = 1, \tag{3}$$

$$E\left[K_j^* \frac{P^*}{P_j^*} \frac{S}{S_j} R\right] = 1. \tag{4}$$

These equations can be used to derive a relation between the currency prices and country pricing kernels.

The following no-arbitrage analysis is a simplified version of Backus et al. (1996). From Eqs. (1) and (4), we have that

$$\Sigma_j \pi_j \left[K_j^* \frac{P^*}{P_j^*} \frac{S}{S_j} - K_j \frac{P}{P_j}\right] = 0, \tag{5}$$

while Eqs. (2) and (3) imply

$$\Sigma_j \pi_j \left[K_j^* \frac{P^*}{P_j^*} - K_j \frac{P}{P_j} \frac{S_j}{S}\right] = 0. \tag{6}$$

How can Eqs. (5) and (6) both be satisfied simultaneously? It is easy to see that a sufficient condition is:

$$\frac{K_j^*}{K_j} = \frac{P}{P_j} \frac{P_j^* S_j}{P^* S}, \qquad \text{for all } j. \tag{7}$$

If markets are complete, this is also necessarily the case (with $j=1$, 2, Eqs. (5) and (6) yield two linear equations in K_1^*/K_1 and K_2^*/K_2). If the set of markets is incomplete, the K_j^*/K_j ratios can nevertheless still be chosen such that Eqs. (5) and (6) hold and that no-arbitrage opportunities exist.

We rewrite the no-arbitrage relation (7) in log-form

$$\Delta s = (\Delta p - \Delta p^*) + k^* - k, \tag{8}$$

where $\Delta x = \log(X_j/X)$. Combine Eq. (8) with the quantity equation $MV = PY$. Rework this definition into the logarithmic relative country version of the quantity equations in first differences

$$\Delta p - \Delta p^* = (\Delta m - \Delta m^*) + (\Delta \upsilon - \Delta \upsilon^*) - (\Delta y - \Delta y^*). \tag{9}$$

Define a relative country variable as $\Delta \tilde{x} = \Delta x - \Delta x^*$. Combine Eqs. (8) and (9) to obtain an expression for the forex returns:

$$\Delta s = \Delta \tilde{m} - \Delta \tilde{y} + \Delta \tilde{\upsilon} - \tilde{k}. \tag{10}$$

Note that up to this point, the expression for the exchange rate is derived from first principles and should therefore apply to any exchange rate regime. The more conventional derivation starts from Eq. (9) and imposes (relative) PPP, which is a somewhat more specific no-arbitrage assumption. This directly yields $\Delta s = \Delta \tilde{m} - \Delta \tilde{y} + \Delta \tilde{\upsilon}$, but one looses the insight that deviations (from PPP) perturb the relative intertemporal marginal rates of substitution $\tilde{k}$ in a specific way. To make the concept of velocity υ operational, however, we still have to take a shortcut. Present-day monetary theory does not yield a standard structural velocity specification.[1] For this reason, we take the direct macro approach and assume that velocity is a stable function of the interest rate and income. This is the Achilles heel to our analysis of regime change. In the end, though, we can relax the stability of the country velocity specifications somewhat, since we only need the relative country variables. The log-velocity specification is then as follows:

$$\upsilon = \lambda(R-1) + \tau y + \varepsilon,$$

and in relative country variable notation:

$$\tilde{\upsilon} = \lambda \tilde{R} + \tau \tilde{y} + \tilde{\varepsilon}, \tag{11}$$

where $\tilde{\varepsilon}$ is the unexplained part of the relative velocity.

[1] Standard theoretical approaches are the monetary overlapping generations model, a transactions technology approach whereby money enters the utility function indirectly via leisure time, the cash-in-advance constraint-based analysis, and the decentralized exchange model. But neither approach generates a canonical expression for velocity.

Substitute Eq. (11) in Eq. (10) to arrive at

$$\Delta s = \Delta \tilde{m} + (\tau - 1)\Delta \tilde{y} + \lambda \Delta \tilde{R} - \tilde{k} + \Delta \tilde{\varepsilon}. \tag{12}$$

The levels version of Eq. (12) is amenable to regression analysis. The last two terms $\Delta \tilde{\varepsilon} - \tilde{k}$ in Eq. (12) are the unobserved residual. In the empirical section, a panel estimation procedure will be used to estimate the coefficients $\tau - 1$ and λ. In the rest of this section, we assume that the homogeneity property of s with respect to $\tilde{m}$ is satisfied, but the empirical analysis does not impose this neutrality restriction. The coefficients are, however, restricted to be identical across countries. The theoretical reason for this restriction is that the structural model is not country-specific. The arguments on which the above specification is built apply to any pair of countries. The cross-country coefficient restriction is also instrumental for the empirical analysis, given the limited amount of data in our panel of yearly observations.[2] But we do investigate whether the restriction on the cointegrating vector is implicitly validated by the data through testing for stationarity of the residuals.

Once we know the coefficients, we can investigate the implications of a monetary union. To this end, rearrange Eq. (12) to obtain:

$$\Delta \tilde{m} = \Delta s + (1 - \tau)\Delta \tilde{y} - \lambda \Delta \tilde{R} + u, \tag{13}$$

where $u = \tilde{k} - \Delta \tilde{\varepsilon}$. To describe the endogenous response of money demand to monetary unification, we need to analyze how $\mathrm{Var}[\Delta \tilde{m}]$ is affected by monetary unification. Evidently, we have the following pre-union decomposition:

$$\begin{aligned} \mathrm{Var}[\Delta \tilde{m}] = {} & \mathrm{Var}[\Delta s - \lambda \Delta \tilde{R}] + (1 - \tau)^2 \mathrm{Var}[\Delta \tilde{y}] \\ & + 2\mathrm{Cov}\left[\Delta s - \lambda \Delta \tilde{R}, (1 - \tau)\Delta \tilde{y}\right] \\ & + 2\mathrm{Cov}\left[\Delta s - \lambda \Delta \tilde{R} + (1 - \tau)\Delta \tilde{y} + u, u\right]. \end{aligned} \tag{14}$$

Given the panel estimates for the coefficients, estimates for all the terms in Eq. (14) can be constructed from the observed m, s, R and y variables. The last component on the RHS is obtained as a residual by taking the difference between $\mathrm{Var}[\Delta \tilde{m}]$ and the first three components on the RHS.

In the empirical section, we will show that the main component of $\mathrm{Var}[\Delta \tilde{m}]$ is $\mathrm{Var}[\Delta s]$. The variation in $\Delta \tilde{R}$ and $\Delta \tilde{y}$ are orders of magnitude smaller than the $\mathrm{Var}[\Delta s]$. This is essentially the news dominance feature of foreign exchange rate movements. Having established this, what does it imply for the monetary union?

[2] The Groen and Kleibergen (1999) framework allows one to deduce how many different cointegrating vectors are present within the panel. Their test results provide strong empirical evidence for the overall cross-country restriction.

Evidently, $\mathrm{Var}[\Delta s]=0$ under monetary union, and $\mathrm{Var}[\Delta \tilde{R}]$ will be approximately equal to 0. Hence, Eq. (14) reduces to:

$$\mathrm{Var}[\Delta \tilde{m}] = (1-\tau)^2 \mathrm{Var}[\Delta \tilde{y}] + 2\mathrm{Cov}[(1-\tau)\Delta \tilde{y} + u, u] \tag{15}$$

This variance is likely to be much smaller than $\mathrm{Var}[\Delta \tilde{m}]$ over the pre-union situation as given in Eq. (14), because $\mathrm{Var}[\Delta s]$ has disappeared from the RHS in Eq. (15). This is the endogenous reaction of monetary aggregates to EMU.

Regarding $\mathrm{Var}[\Delta \tilde{y}]$, Artis and Zhang (1992) and Fatas (1997) conclude that countries participating in the exchange rate mechanism of the EMS show a stronger business cycle synchronization than countries which did not participate. This suggests that under monetary union $\mathrm{Var}[\Delta \tilde{y}]$ would be reduced still further. Though this would alleviate the problem of asymmetrical real shocks, one of the big problems of the EMU, it also introduces the Hayekian type problem that the centralization of monetary policy will enhance the amplitude of the European business cycle. The likely reduction in $\mathrm{Var}[\Delta \tilde{y}]$ thus strengthens the conclusion that $\mathrm{Var}[\Delta \tilde{m}]$ has to come down.

The conclusion that $\mathrm{Var}[\Delta \tilde{m}]$ is lower under EMU also hinges on a number of empirical assessments about the magnitude of the relationships between the variables. We discuss some statistical issues concerning these empirical assessments. Suppose we can assume, if anything, that the differences in real growth rates do not increase, and that the fluctuations in u do not increase either, then $\mathrm{Var}[\Delta \tilde{m}]$ must come down. Recall that $u = \tilde{k} - \Delta \tilde{\varepsilon}$, where $\tilde{k}$ stands for the relative country difference in the intertemporal marginal rate of substitution, and $\Delta \tilde{\varepsilon}$ is the change in the relative country difference in velocity in so far this is not captured by $\lambda \Delta \tilde{R} + \tau \Delta \tilde{y}$. A priori it seems unlikely that unification increases the variability of the real variable $\tilde{k}$. Neither is it likely that velocity differences between countries vary more due to the unification. These assumptions concerning $\tilde{k}$ and $\Delta \tilde{\varepsilon}$ are our maintained hypotheses and go untested in the analysis below. Thus, assuming that u does not thwart the structural invariance we are seeking, it follows that unification lowers the variability of the country differential money growth rates.

There are some econometric concerns which warrant treatment. Note that u consists of two components, i.e. it is the difference between the numeraire country variable $k - \Delta \varepsilon$ and the other country variable $k^* - \Delta \varepsilon^*$. Because the numeraire country residual shows up in all equations of the panel, the cross-country residuals are correlated. This problem is salvaged through the inclusion of time dummies, which remove year specific means from the data. Consequently, the time-specific effect reduces the cross-country correlation of the residuals by a factor of $1/(n-1)$, where n is the number of countries.

An important econometric issue is the non-stationarity of some of the variables. Typically, s, $\tilde{m}$, and $\tilde{y}$ are non-stationary variables, while the first differences of these variables are stationary. In contrast, the interest differential $\tilde{R}$ is stationary, though autocorrelated. Due to the non-stationarity of the variables s, $\tilde{m}$, and $\tilde{y}$, the difference specification in Eq. (12) is not directly amenable to regression analysis. A simple level regression of s on $\tilde{m}$, $\tilde{y}$ and $\tilde{R}$, suffices to back out consistent coefficient estimates, albeit with different rates of convergence. The short-run dynamics and the long-run equilibrium can be jointly modelled through an error-correction specification. But the short-run dynamics are not of immediate concern for our analysis, since we only need the long-run coefficients for Eq. (14), and hence a level regression procedure suffices.

Given our interest in the long-run coefficients, though, we also want to report some tests of hypotheses concerning these coefficients. A static level regression of s on $\tilde{m}$, $\tilde{y}$ and $\tilde{R}$, will produce consistent coefficient estimates, but does not necessarily provide appropriate standard errors. Standard errors from the level panel regression cannot be used in standard testing procedures if the innovations to $\Delta\tilde{m}$ and $\Delta\tilde{y}$ are correlated with the residual of the levels regression of s on $\tilde{m}$, $\tilde{y}$ and $\tilde{R}$ due to the short-run dynamics. The Stock and Watson (1993) dynamic OLS procedure remedies this problem. The idea is to expand the regression with lags and leads of the first differences of the explanatory variables. Recently, Mark and Sul (1999) showed explicitly that the procedure can be adapted to the panel cointegration setting. This procedure is used in the empirical section to enable tests of hypotheses concerning the long-run coefficients. We also test for cointegration in different ways; see Section 3. The limited number of yearly data, however, imply that these tests have to be interpreted with some caution. Moreover, we investigate the properties of the error correction coefficient across countries assuming stationarity.

Lastly, we touch upon the issues of omitted variable and simultaneity bias. The usage of a panel may, to some extent, redress these curses. This mitigation occurs if the correlation between residual and explanatory variables is policy dependent and hence is different across countries. In fact, the differences in time-series versus cross-section estimates of the coefficients point in this direction. Another reason why a potential bias in the estimate of, e.g. λ may not overly concern us, is that Var[$\Delta\tilde{R}$] is relatively small compared to Var[Δs]. Lastly, if the relation between the omitted variable and the included variable is stable across the regime change, then the omitted variable bias is a virtue since it captures both the effect of how the dependent variable changes due to changes in the dependent variable, and the effect of the concomitant changes in the omitted variable.

With these caveats in mind, we can now explore the economic implications of the predicted decline in Var[$\Delta\tilde{m}$] due to unification. By definition, the following decomposition holds:

$$\mathrm{Var}[\Delta\tilde{m}] = \mathrm{Var}[\Delta m] + \mathrm{Var}[\Delta m^{*}] - 2\mathrm{Cov}[\Delta m, \Delta m^{*}]. \tag{16}$$

Our prediction is that Var[$\Delta\tilde{m}$] will decline substantially, as the variability of all three independent variables in Eq. (14) will be reduced. Suppose that Var[Δm] and Var[Δm^{*}] settle somewhere in the neighborhood of their pre-union average (Var[Δm] + Var[Δm^{*}])/2, say, on the grounds that the current union's monetary policy stance is somewhere in between the pre-union positions of the member countries (note that this is only one of the possible scenarios). Conditional on this assumption, the pre- and post-union sum $\mathrm{Var}[\Delta m] + \mathrm{Var}[\Delta m^*]$ are about equal. Therefore, the covariance $\mathrm{Cov}[\Delta m, \Delta m^*]$ has to go up. This implies an increase in the comovement of money across nations. The corollary to this effect is that the apparent law of large numbers effect for the pre-union monetary aggregates must be absent in the post-union data, as the correlations between the Δm_i will tend to 1; the LHS in Eq. (16) will become very small, and $\mathrm{Var}[\Delta m] \approx \mathrm{Var}[\Delta m^*]$. Hence, our claim that the averaging effect is illusory, and that rather the reverse can be expected, as the empirical evidence by De Grauwe (1996) and Arnold (1997) shows. Our exchange rate model predicts that all Δm_i behave very similarly due to unification.

As a rough approximation to the unification effect, we can set the LHS in Eq. (16), i.e. $\mathrm{Var}[\Delta \tilde{m}]$, equal to 0. This more or less implies that the Δm and Δm^* become perfectly correlated, i.e. the endogenous response is complete. Separate from this is the level of the variability of Δm and Δm^* that is experienced in the monetary union. Regardless unification, the following expression for Δm holds:

$$\Delta m = \Delta p + \Delta y - \Delta v.$$

It is well known that most of the variation in Δm ends up in Δp; to quote Friedman (1963): 'Inflation is always and everywhere a monetary phenomenon'. Hence, the level of inflation variability experienced depends on the monetary prudence exercised by the ECB. The quality of ECB policy is, however exogenous to our model and does not depend on the endogenous response that we predict. The endogenity in European money demand will be present regardless the level of prudence exercised by the ECB, so we can leave aside the predictions about ECB policy. This also implies that there is no free lunch for the ECB from the unification, as the above quoted studies have suggested on the basis of the variance of the average effect. Monetary prudence has to be gained the hard way, by proper non-activist monetary policy. In any case, European money demand studies using constructed monetary aggregates under the old regime are of little use, due to the change of regime and the ensuing endogenous response.

3. Empirical validation

For the empirical implementation, we consider the long-term relationship between the variables. Short-run relations between the instruments of monetary policy and monetary targets such as inflation are rather opaque. The relations depend on institutional details and specific circumstances that differ across countries and that are difficult to model, but monetarists maintain that the equilibrium transmission mechanism is stable and fairly transparent. For this reason, central bankers of stable monetary regimes tend to be unperturbed by short-run deviations between monetary targets and realizations, witness the fact that the Bundesbank on many occasions missed its annual money growth targets. Instead, policy decisions are based on medium-run developments, exploiting the known long-run equilibrating responses between the different variables. The knife thus cuts both ways. For several reasons, we do not know much about the short-run relations between the various monetary variables, and hence policy bases itself on longer-term developments. Therefore, from a EURO monetary policy point of view, money demand stability will not be an issue of concern on a day-by-day basis. This seems also to be the stance that is practised by the ECB.

Within the setting of the foreign exchange market, there are two specific arguments for focussing on the medium- to longer-term relations between the variables. First, under a free float regime, the spot exchange rate incorporates the expectations regarding future deviations between domestic and foreign money growth. Hence, the relation between the relative money supply and the spot rate at any point in time can be rather loose, especially for the econometrician who does not observe these expectations. Empirically, Meese and Rogoff (1983) demonstrated that the fundamentals-based foreign exchange rate model has no superior forecasting power over the simple no change forecast derived from the martingale feature of asset market pricing. Turning this evidence around, one could say that as an equilibrium relation the fundamentals-based models, like the levels version of Eq. (12), is not rejected, but also that it contains no information about how the exchange rate will move during the next specific instant. Second, in fixed or managed exchange rate systems, like the ERM, exchange rate adjustments reflect a divergence of money growth and inflation that developed in the past. Because governments have committed to certain exchange rate targets, it is public knowledge that adverse movements in the underlying variables have to be countered sooner or later to satisfy the long-run equilibrium relation between the fundamentals and the exchange rate. This induces the stationary autoregressive behavior between the set of fundamentals that is so typical for the target zone models. For example, the German interventions on behalf of the French Franc after the demise of the Lira and the Pound in 1992 induced short-term volatility in the money figures. The interventions, however, did not lead to increased exchange rate variability, because the interventions countered the expected changes in the exchange rates, and these therefore did not materialize at that time.[3] But if no corrective action is undertaken on the level of the fundamentals, then eventually exchange rate adjustments will do the job. For both reasons, we do not expect our

[3] We note that such sizable shifts between ‘nickels and dimes’ has no significance for the case of monetary union, except perhaps during the transition phase when exchange rates have been irrevocably fixed, but national denominations still circulate.

model to be applicable in the short-run. We do expect, however, that systematic deviations between domestic and foreign money growth will either be foreshadowed in the forex markets or be corrected through realignments, so that Eq. (12) in levels still holds in the long run.

The panel procedure is especially suited to capture the long-run equilibrium relation, because short-run deviations, which differ across countries depending on foreign exchange rate regimes, are averaged out across countries. A time-series analysis for a specific exchange rate is therefore less well suited for finding the long-run coefficients, because a particular country may stay on the same regime during the entire sample. Apart from having to estimate the coefficients of the exchange rate Eq. (12), we also need the various second moments for Eq. (14) per pairs of countries. Just as the time-series analysis is not simple, it is not easy to disentangle the variability in the persistent innovations from the variability in the short-run disturbances for the relative country variables. It is a nonetheless straightforward exercise to show that the identification of the persistent factors versus the transitory components in the disturbances is facilitated by computing the variances over longer horizons. For this reason, we compute the variances and covariances using a multiperiod horizon, so that the transitory factors only play a minor role.

Some intuition for the results to come can be given graphically. Fig. 1 plots time series for the income velocity of money for 14 European countries and eight US regions over the period 1974 to 1988.[4] The velocities of the nine European countries plotted in Fig. 1a wander in all directions. In contrast, Fig. 1b suggests that the velocities of the five 'core' countries Austria, Belgium, France, Germany and The Netherlands, share a common downward trend. Finally, Fig. 1c shows that within the United States, even short-run fluctuations are to a large extent synchronized, witness the sharp drop in velocity in 1986. This suggests that as countries or regions move closer towards monetary union, the comovement in velocities endogenously increases.

Before we can proceed with the variance decomposition based on Eq. (14), we need to estimate the parameters of the monetary exchange rate model. We use annual data covering the period 1975 to 1995. The data were take from European Economy. The sample consists of all countries of the European Union, except Luxemburg, but including The United States and Japan, i.e. 15 countries in total. For each country, we have taken the relevant broad monetary aggregate, which is either M2, M3 or M4 (for Britain). Real income is measured by GDP. We use the

[4] The scale difference between the European and the US series is due to the choice of monetary aggregate. For the European countries, we use a broad monetary aggregate, the US series are based on demand deposits only. We are interested in the coherence, not the scale. The sample period is constrained by the availability of the US data.

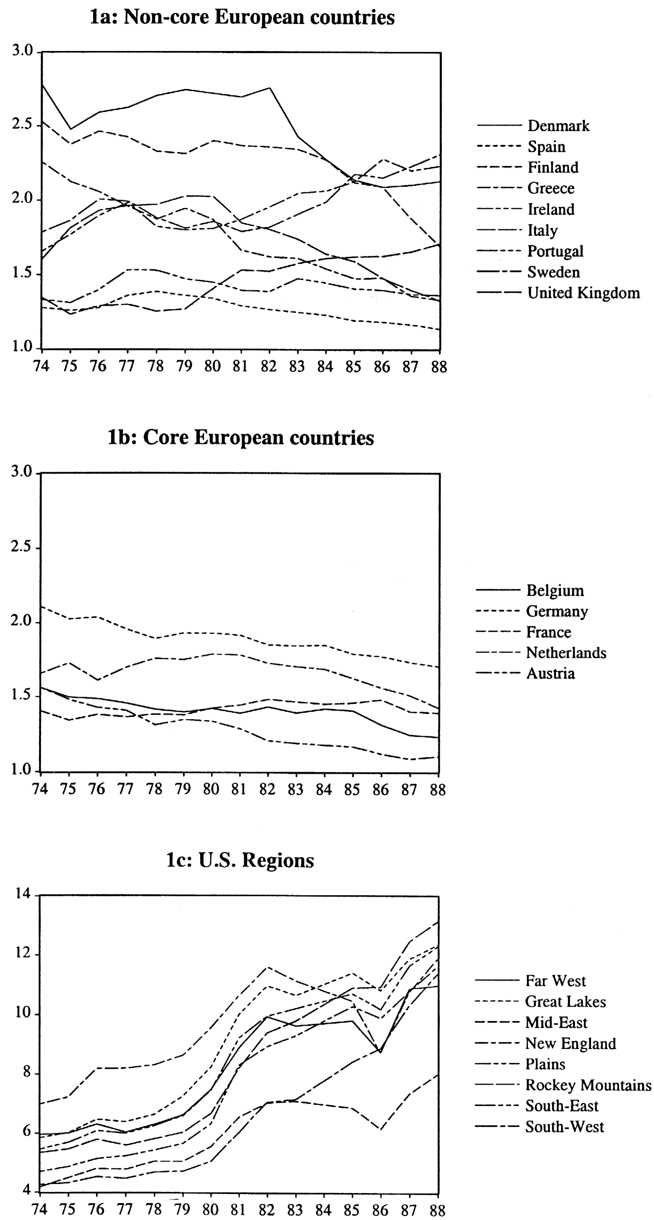

Fig. 1. Coherence in velocity. (a) Non-core European countries. (b) Core European countries. (c) US Regions.

short-term 3 months interest rate to construct the interest differential (comparable yearly rates were not sufficiently well available). Missing interest rate data for Sweden, Spain and Greece at the beginning of the sample were 'constructed' by using the relationship between the discount rate, which was available over the whole period, and the short-term interest rate. The alternative solution, to delete these countries from our sample, hardly affects the following results. Exchange rates are end-of-year quotes. Data for s, $\tilde{m}$ and $\tilde{y}$ are taken in deviation from their mean in the panel estimation procedure, to account for scale differences between the country variables.

The panel dynamic OLS estimation consists of a levels regression between s, $\tilde{m}$, $\tilde{y}$, $\tilde{R}$, the time dummies d_t, and the one-period leads and lags of $\tilde{m}$, $\tilde{y}$, $\tilde{R}$, per country vis-a-vis the benchmark country, for the model from the previous section:

$$\begin{aligned} s_{it} = {} & c + \beta\tilde{m}_{it} + (\tau - 1)\tilde{y}_{it} + \lambda\tilde{R}_{it} + d_t + a_{1i}\Delta\tilde{m}_{it+1} + a_{2i}\Delta\tilde{y}_{it+1} \\ & + a_{3i}\Delta\tilde{R}_{it+1} + a_{4i}\Delta\tilde{m}_{it-1} + a_{5i}\Delta\tilde{y}_{it-1} + a_{6i}\Delta\tilde{R}_{it-1} + e_{it}, \\ & i = 1,\ldots,n, \quad \text{and} \quad t = 1,\ldots,T, \end{aligned} \tag{17}$$

where $n+1$ is the number of countries, and e_{it} denotes the error term. Table 1 gives a summary report of the estimation results for the relevant coefficients. In total, there are 90 coefficient estimates for the leads and lags, some significantly different from zero, but since our interest lies in the long-run coefficients we do not report these (the estimates are available upon request). All the three coefficient estimates do have the anticipated sign, and are significantly different from zero. The estimate of β conforms surprisingly well with the monetary exchange rate model, i.e. the exchange rate is close to being linearly homogeneous in the relative quantity of money, but the t-test rejects exact homogeneity (t-value of 4.02). Univariate time-series analyses report a wide diversity of β estimates, mostly far from the theoretical value of 1. Because the panel incorporates a large number of countries with quite different monetary policy regimes, the panel is much more informative about the long-run equilibrium relation (17). The estimated coeffi-

Table 1
Panel dynamic OLS estimates

	Eq. (17)	
	Coefficient	Standard error
c	−0.01	0.04
β	0.81	0.05
$(\tau - 1)$	−1.72	0.22
λ	1.46	0.35

cients are robust to omitting individual countries and years from the sample.[5] We also note that the parameter estimates are invariant under the change of numeraire, for example if the US were used instead of Germany.

In testing for cointegration in a panel, one plausible hypothesis is that the cointegrating vector applies to all countries. It is however less clear what the other hypothesis should maintain. Suppose that, counter to actual practice, a group of countries had decided to coordinate their respective money supplies such that the relative money supply variables $\tilde{m}_{it}$ became stationary; in that case the s_{it} would still be cointegrated with $\tilde{y}_{it}$ but not with $\tilde{m}_{it}$, while for the other countries $\tilde{m}_{jt}$ would also be part of the cointegrating vector. The problem is that the alternative is a composite hypothesis of many conceivable economically relevant alternatives. We selected three sets of hypotheses that we deemed interesting.

The first is a test of the null hypothesis that the residuals from the panel dynamic OLS are driven by a non-stationary process with identical coefficients across countries, versus the alternative that the residuals for all countries are driven by a stationary process, again with identical coefficients across countries. The second test relaxes the restriction that the coefficient on the stationary residuals under the alternative hypothesis, i.e. the error correction coefficient, is identical across countries. The third hypothesis supposes that all countries have the same error correction mechanism, versus the alternative that the error correction coefficients differ per country (but both hypotheses maintain that the residuals are stationary). In a way, we regard this test as more interesting than testing for cointegration. For sure, the third test takes cointegration as given, but economic theory is quite adamant that this should be the case. We note that while it is easy to formulate the null of no cointegration statistically, there are few if any theoretic exchange rate models that would be in conformity with this hypothesis. It is, however, not difficult to come up with economic models where the adjustment coefficients do differ across countries, due for example to differences in monetary policy operating procedures. Hence, it is certainly of economic interest to know whether or not the error correction coefficients differ per country under the maintained hypothesis of stationarity.

For the first test, we use the results in the upper part of Table 2, which are based on a pooled ADF regression on the estimated residuals from the dynamic panel OLS regression. For the choice of critical value, we assume that the number of countries is fixed, but that time can expanded indefinitely. It is also assumed that the system is driven by just three non-stationary variables, and by a number of

[5] In our working paper version (Arnold and de Vries, 1998), we report the coefficient estimates based on the level regression without the leads and lags of the first differenced variables (first step from the two-step cointegration procedure), as 0.85, -1.70 and 0.50, for β, $\tau-1$ and λ, respectively. While different, the ensuing analysis is hardly affected.

Table 2
Cointegration tests

	Coefficient	Standard error	t-value
overall: $\Delta e_{it} = \gamma e_{it-1} + \zeta \Delta e_{it-1} + \eta \Delta e_{it-2} + u_{it}$			
γ	−0.46	0.07	−6.12
ζ	−0.29	0.07	−3.82
η	−0.11	0.06	−1.74
	# observations 240, $R^2 = 0.37$		
per country: $\Delta e_{it} = \gamma_i e_{it-1} + u_{it}$			
γ_{BE}	−0.44	0.19	−2.28
γ_{DK}	−0.77	0.24	−3.23
γ_{ES}	−0.65	0.24	−2.74
γ_{FI}	−0.54	0.26	−2.10
γ_{FR}	−0.41	0.19	−2.11
γ_{GR}	−0.70	0.23	−3.05
γ_{IR}	−0.40	0.19	−2.09
γ_{IT}	−0.69	0.28	−2.44
γ_{NL}	−0.50	0.21	−2.38
γ_{AU}	−0.27	0.15	−1.74
γ_{PO}	−1.08	0.23	−4.75
γ_{SE}	−0.91	0.22	−4.07
γ_{UK}	−0.52	0.17	−3.03
γ_{US}	−0.83	0.24	−3.48
γ_{JP}	−0.36	0.17	−2.13

stationary variables. Correlation between the residuals across countries causes size distortions (see Groen, 2000). Suppose the residuals are perfectly correlated across countries, but not the other right hand side variables. This essentially reduces the number of observations in the cointegration test regression to the number of time periods. To stay on the safe side, we therefore obtain the MacKinnon (1991) critical value by calculating the correction factors and using the number of observations per country 16, rather than the total number 240. This results in a critical value equal to −3.75 (if the number 240 is used the critical value becomes −3.36), while our test statistic equals −6.12. Hence, on basis of this, we reject the null in favor of the hypothesis of cointegration.

Under the second test, the null hypothesis is unaltered, but the alternative hypothesis allows for different error correction coefficients. These coefficients can be estimated by running the Dickey–Fuller regressions per country; see the lower part of Table 2 for the results. Because the observations per country are so few relative to the entire sample, and since the dynamic OLS estimates for the coefficients on the non-stationary variables are super consistent, a first cut would be to use the Dickey–Fuller critical value (−1.96). In that case, 14 out of the fifteen γ_i estimates are significantly below 0, leading to a rejection of the null. A more conservative test would be to use the MacKinnon (1991) critical value. If the residuals are independent across countries, then one would expect to see about one out of the 15 ($1/15 = 0.066$) t-statistics below the 5% critical value -3.75 under the null of no cointegration. We find two t-values below -4, casting doubt on the null hypothesis but not leading to a sound rejection.[6]

The third test investigates whether all countries have the same error correction mechanism or not, assuming cointegration. Suppose the null is that the error correction coefficient equals -0.46, the value of γ we find from the panel cointegration regression in the upper part of Table 2. From the standard errors in the lower part of the Table 2, we see that only one estimate, γ_{PO}, is significantly below the 5%-critical value, which is as expected under the null in a 'random' sample of fifteen countries. This means that all countries in the sample respond in a remarkably similar fashion to deviations from the long run equilibrium (this is not withstanding the fact that the short run dynamics can still be different). The evidence thus provides little support for the often voiced concern that the ECB has to take into account differences in transmission across countries when formulating its monetary policy.

The estimates for β, τ, and λ are subsequently used for a variance decomposition based on Eq. (14). As we argued above, one needs to worry about the 'long and variable lags'. In order to identify the persistent factors and keep the contribution of the transitory elements to a minimum, we base the calculations of the second moments on non-overlapping 4-year period growth rates for all the variables concerned. For our sample period, this leaves five observations per exchange rate. Nevertheless, this loss in precision is acceptable, given that so many countries are present in the sample. Table 3 reports for all rates the following terms in the decomposition: $\beta^2 \text{Var}[\Delta \tilde{m}]$, $\text{Var}[\Delta s - \lambda \Delta \tilde{R}]$, $(1-\tau)^2 \text{Var}[\Delta \tilde{y}]$, and $2\text{Cov}[\Delta s - \lambda \Delta \tilde{R}, (1-\tau)\Delta \tilde{y}]$. The numbers for the last element $2\text{Cov}[\Delta s - \lambda \Delta \tilde{R} + (1-\tau)\Delta \tilde{y} + u,\ u]$ on the RHS of Eq. (14) are constructed from the other elements. Table 3 also reports Spearman's rank correlations r_S between column (a) and the columns (b), (c), (d) and (e); t-statistics t_{r_S} are in the last row. The final column (f) gives the 1995 value of M/M^{GE} indexed with base year 1974, as an indicator of long-run monetary divergence. The rank correlation r_S is in this case between columns (b) and (f).

The results show that the three countries with lowest exchange rate variability vis-a-vis Germany, Austria, Belgium, and The Netherlands — also have the lowest variability in the money growth differential. In contrast Italy, Greece and

[6] Note that one cannot use the individual country tests to conclude that the countries with t-statistics below the critical values are the countries for which there is cointegration, while for the other countries there is no cointegration; such a conclusion necessitates that the null identifies apriori what one expects to be the case for each country.

Table 3
Variance decomposition

	(a)	(b)	(c)	(d)	(e)	(f)
BE	26	82	23	64	−142	129
DK	176	102	113	−52	13	173
ES	143	405	48	−126	−184	436
FI	228	424	218	−313	−101	239
FR	89	184	28	43	−165	153
GR	517	781	37	64	−365	1504
IR	194	101	109	−42	25	297
IT	264	574	54	205	−570	279
NL	21	10	23	−19	7	103
AU	33	5	21	−10	16	139
PO	560	2959	24	372	−2795	932
SE	67	151	107	43	−235	121
UK	164	179	118	111	−245	309
US	65	1467	114	188	−1705	107
JP	87	519	62	−188	−306	152
r_S with (a)		0.61*	0.33	0.11	−0.34	0.45**,a
t_{r_S}		2.78	1.27	0.68	1.30	1.83

Column (a) — $\beta^2 \text{Var}[\Delta \tilde{m}]$.
Column (b) — $\text{Var}[\Delta s - \lambda \Delta \tilde{R}]$.
Column (c) — $(1-\tau)^2 \text{Var}[\Delta \tilde{y}]$.
Column (d) — $2\text{Cov}[\Delta s - \lambda \Delta \tilde{R}, (1-\tau)\Delta \tilde{y}]$.
Column (e) — $2\text{Cov}[\Delta s - \lambda \Delta \tilde{R} + (1-\tau)\Delta \tilde{y} + u, u]$.
Column (f) — index(M/M^*).

*Significant at a 5% level.

**Significant at a 10% level.

[a] r_S with (b), all others with (a).

Portugal combine high exchange rate variability with high variability in the money growth differential. The combination of high exchange rate variability and low variability in the money growth differential in the United States and Japan does not invalidate our argument, since we only argue that a lack of monetary comovement will be reflected in the exchange rate, not that real exchange rate fluctuations are impossible. Taking all countries together, Table 3 shows that $\text{Var}[\Delta s - \lambda \Delta \tilde{R}]$ is the only component in the variance decomposition that is significantly related to $\beta^2 \text{Var}[\Delta \tilde{m}]$, as indicated by the rank correlations. The last column in Table 3 shows that our other indicator of (a lack of) monetary comovement is significantly related to $\text{Var}[\Delta s - \lambda \Delta \tilde{R}]$ at a 10% level.

In addition to the results for the individual countries, we investigated three country groupings: A small group, consisting of Austria, Belgium, France and The Netherlands; a medium-sized group also including Finland, Ireland, Italy, Portugal and Spain; and a large group including all EU-members (with the exception of Luxemburg). The small group consists of countries with a track record of

Table 4
Average EMU variance decomposition

	(a)	(b)	(c)	(d)	(e)	(f)
Small EMU	42	70	23	19	−71	131
EMU	173	527	61	19	−434	301
Large EMU	191	458	71	26	−365	370

Column (a) — $\beta^2 \mathrm{Var}[\Delta \tilde{m}]$.
Column (b) — $\mathrm{Var}[\Delta s - \lambda \Delta \tilde{R}]$.
Column (c) — $(1-\tau)^2 \mathrm{Var}[\Delta \tilde{y}]$.
Column (d) — $2\mathrm{Cov}[\Delta s - \lambda \Delta \tilde{R}, (1-\tau)\Delta \tilde{y}]$.
Column (e) — $2\mathrm{Cov}[\Delta s - \lambda \Delta \tilde{R} + (1-\tau)\Delta \tilde{y} + u, u]$.
Column (f) — $\mathrm{index}(M/M^*)$.

exchange rate stability vis-a-vis Germany, while the medium-sized group includes all the EMU-participants. The numbers in Table 4 are averages of the numbers for the individual countries from Table 3. The results for the country groupings confirm the difference between the 'core' countries and the other European countries that we observed in Fig. 1. Table 4 shows a striking difference between the small group and the other two groupings: In the small group, the variability of $\Delta s - \lambda \Delta \tilde{R}$, $\Delta \tilde{m}$ and $\Delta \tilde{y}$ is much lower than in the other groups.

What endogenous response will EMU deliver? Due to the unification, $\mathrm{Var}[\Delta s - \lambda \Delta \tilde{R}]$ and $\mathrm{Cov}[\Delta s - \lambda \Delta \tilde{R}, (1-\tau)\Delta \tilde{y}]$ drop out. But since $\Delta s - \lambda \Delta \tilde{R}$ is also part of $\mathrm{Cov}[\Delta s - \lambda \Delta \tilde{R} + (1-\tau)\Delta \tilde{y} + u, u]$, we assume that this covariance drops out as well, to stay on the safe side (note that the column (e) entries in both tables are mostly negative; so that retaining a part of these negative covariances would only reinforce our conclusions). Thus, our post-union Eq. (15) is reduced to $\mathrm{Var}[\Delta \tilde{m}] = (1-\tau)^2 \mathrm{Var}[\Delta \tilde{y}]$. Hence, we can focus on columns (a) and (c) to distill the endogenous response to monetary union. Consider the case of a small EMU. From Table 4, we find that for the small EMU, $\beta^2 \mathrm{Var}[\Delta \tilde{m}]$ goes from 42 in the pre-union situation to 23 post-union, a decrease of nearly 50%. This increases the comovement between the local inflation rates by the same amount if the average of the variances remains about constant, cf. Eq. (16). For the EMU and large EMU variants, the predicted increase is much larger. See also column (c) in Table 3 for the individual countries.[7] Note that these predictions are conditional on the past level of $(1-\tau)^2 \mathrm{Var}[\Delta \tilde{y}]$. If EMU leads to stronger business cycle synchronization, $\beta^2 \mathrm{Var}[\Delta \tilde{m}]$ goes down even more.

[7] Sweden is the only EU country for which $(1-\tau)^2 \mathrm{Var}[\Delta \tilde{y}]$ exceeds $\beta^2 \mathrm{Var}[\Delta \tilde{m}]$ by a large amount.

4. Conclusions

A truly European monetary policy can only exist if monetary impulses from the ECB are transmitted rapidly and uniformly throughout the eurosystem. This requires complete money market integration. Full money market integration induces a comovement among the euro-denominated monetary aggregates of the EMU countries, as changes in monetary policy and money demand affect all the aggregates simultaneously, whatever course they take. This endogenous response to unification invalidates European money demand studies which have a built-in bias towards stability due to the law of large numbers. Our evidence indicates that this bias is least severe for European money demand studies which include only those countries, which have effectively operated as a monetary union. For example, a study of German–Dutch money demand from March 1983 onwards, when the guilder was last devalued, would not fall victim to our critique (compare the variances 21 and 23 from the columns (a) and (c) in Table 3; the predicted increase in comovement is, in this case, essentially zero). Unfortunately, such a study would have very limited relevance for monetary policy in the EMU. As the recent survey by Browne et al. (1997) shows, most European money demand studies use a large group of countries (always including Italy) and a sample period starting in the 1970s. Since these samples lack exchange rate stability vis-a-vis Germany, we cannot use them to make inferences regarding money demand stability under the EMU regime. Thus, we maintain that the average of pre-union local money demand has no predictive power whatsoever. Rather, a better predictor for things to come is to pick the individual country whose past institutional central bank design best reflects the current design of the ECB, and to extrapolate from there. We leave this exercise to the reader.

The endogeneity in European money demand described in this paper is a nice example of what the Lucas critique is all about. The implication for current monetary policy making by the ECB is as follows. Given that there is no law of large numbers effect due to monetary unification, there is also no extra room for a more activist monetary policy stance. Monetary stability is not a bonus awarded to politicians for completion of the long march towards EMU, but has to be earned the hard way, i.e. by prudent monetary policy. One can perhaps understand the calls for a more activist stance from the misperception that arises from the studies based on the average of pre-union money demand functions. But this is not an argument for macro experiments. The added difficulty is that data take a long time to prove, which argument is correct, due to the long and variable lags in the transmission of monetary policy initiatives; let alone that macro experiments can be very costly. The above combination of theoretical and empirical arguments can hopefully be used to reduce pressure for activism on current monetary policy making.

Since its inception, two other aspects of ECB policy making have also been criticised, apart from the particular monetary policy stance taken by the ECB. The first criticism relates to the ECB's lack of transparency: witness its secretive decision-making and failure to publish minutes of meetings or economic forecasts. But also its two-pillared monetary strategy — based on trends in money growth and a broad assessment of the inflation outlook — has come under attack, see Favero et al. (2000). Some critics have denounced the monetary pillar as a smokescreen or, at best, an ineffectual device to transfer the Bundesbank's reputation to the ECB. In contrast, other commentators have criticized the present neglect of the monetary pillar and favour a return to the Bundesbank's policy of monetary targeting. Our paper shows that, though typical for the compromise decision-making within EU institutions, the ECB's present strategy is not necessarily bad. Until the dust of monetary unification has settled, it would be unwise to pin down ECB policy to money growth targets. The present strategy offers the ECB the flexibility that it needs when economic relationships of the past cannot be relied upon; but, at the same time the above arguments provide sufficient armour against activism. In the future, the ECB may activate its currently dormant monetary pillar, when the endogenous response in European money demand has run its course. In the meantime, it can work on its transparency by publicly dispelling illusory benefits from monetary unification.

Acknowledgements

We benefited from presenting some of this material at the ZEI summer school, Bonn 1997, the international economics seminar at the KU Leuven and at the EEA conference, Berlin 1998. We are grateful to Paul De Grauwe, Harris Dellas, Martin Fase, Jan Groen, Heejon Kang, Patrick Minford, Dave Smant, the referees and the editor Arye Hillman for very helpful comments.

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