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Unit roots and Granger causality in the EMS interest rates: the German Dominance Hypothesis revisited

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Abstract

The aim of this paper is twofold: First, it shows that: (a) sufficient conditions for unit roots, found in AR systems, to persist in VAR systems amount to Granger non-causality in any direction among the variables involved. This implies that a necessary condition for the disappearance of one unit root in a VAR system implies Granger causality in at least one direction; (b) for first-order models with non-explosive variables, Granger causality is also sufficient for cointegration; and (c) causality and cointegration inference are strongly affected by the omission of an important causing variable. Second, our new analytical framework allows us to formulate and test several new variations of the so called German Dominance Hypothesis, in a unified cointegrating framework, which we define as Strong, Semi-Strong and Weak (types 1 and 2) German Dominance. In addition, we allow for the possibility that the dominant player may lie outside the EMS, so that we introduce and test the 'US Dominance Hypothesis'. © 1999 Elsevier Science Ltd. All rights reserved.

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1. Introduction

The symmetric property of the European Monetary System (EMS) was questioned since its establishment in 1979. Since then, the popular view has been that it evolved into a non-symmetric system, the members of which do not equally share the burden of adjustment for eliminating balance of payments disequilibria. According to this view, known as the German Dominance Hypothesis (GDH), this asymmetric system has been dominated by the German Central Bank. That is, the Bundesbank independently chooses its monetary policy, i.e. it fixes the reference level of interest rates and controls the exchange rate of the ECU vis-à-vis the US dollar, while the other EMS national central banks stabilize the parity of their currency vis-à-vis the Deutsche Mark (DM), which in turn determines their own domestic exchange rates. Thus, the national central banks follow the Bundesbank's low inflation policy rule and gain credibility (Giavazzi and Pagano, 1988; Von Hagen and Fratianni, 1990; Artus et al., 1991; Kirchgaessner and Wolters, 1993). Hence, the EMS is a de facto DM-zone.

The arguments made and the evidence cited in support of the GDH are mixed. Several studies have claimed that the national central banks have surrendered their monetary policy autonomy to the hegemony of the Bundesbank (Fischer, 1987; Giavazzi and Giovannini, 1987; Wyplosz, 1989; Artis and Nachane, 1990; Karfakis and Moschos, 1990; Biltoft and Boersch, 1992; Caporale and Pittis, 1993). Nonetheless, this strict German Dominance Hypothesis has been challenged on the grounds that: (a) Germany is not the absolute hegemon of the system, but still is a relatively strong player. This view is the weak version of GDH (Smeets, 1990; Von Hagen and Fratianni, 1990); (b) The monetary policies within the EMS are more symmetric (DeGrauwe, 1989; Katsimbris, 1993; Katsimbris and Miller, 1993); and (c) that US monetary policy may, in some way, dictate German monetary policy (Artus et al., 1991; Katsimbris and Miller, 1993; Kirchgaessner and Wolters, 1993). That latter view, although an interesting one, has not provided indisputable empirical evidence in support of the causality relationships between the US and Germany and/or the other EMS countries. In this regard, Dominguez (1997) emphasizes the importance of the international transmission mechanism and claims that, among the G-3, US monetary policies are the most influential.

Concerning the empirical work, a number of studies have focused exclusively on interest rate linkages within the EMS and aimed at revealing the direction of causation (if any) among these linkages (Karfakis and Moschos, 1990; Katsimbris, 1993; Katsimbris and Miller, 1993). Their research strategies have concentrated on testing for cointegration in the context of a bivariate VAR system, consisting of the German interest rate and the respective rate of each of the other EMS member countries. The null hypothesis under examination of 'no (long run) German dominance' is translated into 'non-cointegration' in the system. These authors acknowledge the fact, that non-cointegration does not necessarily imply non-causality in general and, therefore, they perform standard causality tests in the first-differenced VAR. What they fail to see is that lack of cointegration may simply be the result of a monotonic *convergence* of the member states' rates towards the German rate, which simply results in trending interest rate differentials. In other words, lack of interest rate comovement may merely be the result of gradual interest rate convergence within the sample period under study. [For a thorough discussion on this issue see Caporale and Pittis (1993).]

On the other hand, Weber (1991) has advanced an argument that the EMS is a bipolar system involving a hard currency option supplied by the Bundesbank and a soft currency option supplied by the Banque de France. However, this 'soft' currency option has been challenged on the grounds that it is rather the sluggish response of labor market expectations (due to 'strong' trade unions and 'weak' governments) relative to the fast response of financial markets, when a country joins the EMS, than the adoption of a 'soft' currency alternative of the Banque de France by few non-German EMS countries (Walters, 1990; Baldwin, 1991; Miller and Sutherland, 1991).

The methodologies of the preceding papers in elucidating the evolution of the EMS are varying. They range from the specification and simultaneous estimation of small econometric models for EMS and non-EMS countries (Von Hagen and Fratianni, 1990; Artus et al., 1991) to the devising of indices intended to measure the policy makers' reputation and credibility within the EMS (Weber, 1991). However, the possibility that the true dominant player, may be outside the EMS has not yet been thoroughly investigated. Specifically, the effects of the US monetary policy on each individual EMS country have not been compared with the corresponding effects from German monetary policy in a unified framework allowing monetary policy interactions between the US and Germany.

In this paper, we re-address the issue of testing for causality in bivariate and trivariate VAR systems. In fact, we examine the conditions under which the unit roots, usually found in autoregressive representations of interest rates carry forward to their VAR representations. In the next section, we show that sufficient conditions for the unit roots to persist in VAR systems amount to Granger non-causality in any direction between the variables involved. We also show that a necessary condition for the disappearance of one unit root in the VAR involves Granger causality in at least one way. This can be thought of as a reconfirmation of the Engle–Granger assertion (Engle and Granger, 1987) that at least one way Granger causality is necessary for cointegration. In addition, we demonstrate that in first-order models and for non-explosive time series, causality is also sufficient for cointegration. Finally, we discuss the implications from the omission of an important 'causing' variable on causality and cointegration inference.

In Section 3 we apply the strategy outlined above to the case of the EMS short-term interest rates. Specifically, we examine the causality and cointegration properties of bivariate first-order VAR systems, consisting of the German rate and the respective rate of each individual EMS country. Here, our finding of no cointegration implies that the sufficient condition for causality in at least one direction is not fulfilled. Next, we investigate whether non-cointegrability in the previous systems arise because of the omission of an important 'causing' variable which, in the present context, is chosen to be the US interest rate. Indeed, cointegration seems to characterise all the trivariate systems, which implies, in turn, that causality, between at least two variables in at least one direction, should be present. Testing for the direction of causality is then associated with some interesting hypotheses regarding the interactions of monetary policies among the countries participating in the VAR. In particular, the GDH can take several forms depending on whether and how the US rate causes the rate of each individual EMS country. These cases, are classified as Strong, Semi-Strong and Weak GDH and are analyzed in the context of the trivariate VAR in Section 3. An additional hypothesis, termed as the US Dominance Hypothesis, is also formulated and tested. Section 4 summarizes the main findings and conclusions.

2. Unit roots, Granger causality and cointegration

Let us consider the non-stationary bivariate stochastic process $Z_t = [y_t, x_t]'$ whose autoregressive representation takes the form[1]

$$Z_t = A_0^b + A_1^b Z_{t-1} + E_t^b, \tag{1}$$

or

$$\begin{bmatrix} y_t \\ x_t \end{bmatrix} = \begin{bmatrix} \gamma_0^b \\ \gamma_1^b \end{bmatrix} + \begin{bmatrix} \alpha_{11}^b & \alpha_{12}^b \\ \alpha_{21}^b & \alpha_{21}^b \end{bmatrix} \begin{bmatrix} y_{t-1} \\ x_{t-1} \end{bmatrix} + \begin{bmatrix} e_{1t}^b \\ e_{2t}^b \end{bmatrix}, \tag{2}$$

where e_{1t}^b and e_{2t}^b both are white noises. The elements of the matrix A_1^b are related to the moments of the joint distribution $f(Z_t, Z_{t-1}; \Theta)$ through the following relationships:

$$a_{11}^b = \frac{\mathrm{Cov}(y_t, y_{t-1}) - a_{12}^b \mathrm{Cov}(x_{t-1}, y_{t-1})}{\mathrm{Var}(y_{t-1})}, \tag{3a}$$

$$a_{12}^b = \frac{\mathrm{Cov}(y_t, x_{t-1}) - a_{11}^b \mathrm{Cov}(y_{t-1}, x_{t-1})}{\mathrm{Var}(x_{t-1})}, \tag{3b}$$

$$a_{21}^b = \frac{\mathrm{Cov}(x_t, y_{t-1}) - a_{22}^b \mathrm{Cov}(x_{t-1}, y_{t-1})}{\mathrm{Var}(y_{t-1})}, \tag{3c}$$

$$a_{22}^b = \frac{\mathrm{Cov}(x_t, x_{t-1}) - a_{21}^b \mathrm{Cov}(y_{t-1}, x_{t-1})}{\mathrm{Var}(x_{t-1})}. \tag{3d}$$

[1]To distinguish between the coefficients in bivariate and trivariate systems (Section 2.1 below), we employ superscripts b and r, respectively.

The unit root condition in the bivariate VAR(1) amounts to both eigenvalues of A_1^b being equal to one (see Spanos, 1990):

$$\lambda_1,\lambda_2 = 1 = \frac{a_{11}^b + a_{22}^b \pm \left[(a_{11}^b + a_{22}^b)^2 - 4\det(A_1^b)\right]^{1/2}}{2} \tag{4}$$

or

$$(a_{11}^b + a_{22}^b) - (a_{11}^b a_{22}^b - a_{12}^b a_{21}^b) = 1. \tag{5}$$

A sufficient condition for both eigenvalues to be equal to one, which is equivalent to non-cointegration can take the form;

$$a_{12}^b = a_{21}^b = 0, \tag{6}$$

which is equivalent to saying that there is no Granger causality in any direction between the two variables in the system. This is true, because under Eq. (6) a_{11}^b and a_{22}^b will be equal to:

$$a_{11}^b = \frac{\mathrm{Cov}(y_t,y_{t-1})}{\mathrm{Var}(y_{t-1})}, \tag{7}$$

$$a_{22}^b = \frac{\mathrm{Cov}(x_t,x_{t-1})}{\mathrm{Var}(x_{t-1})}. \tag{8}$$

Then, the non-stationary nature of Z_t implies that $\mathrm{Cov}(y_t,y_{t-1}) = \mathrm{Var}(y_{t-1})$ and $\mathrm{Cov}(x_t,x_{t-1}) = \mathrm{Var}(x_{t-1})$, which in turn results in $a_{11}^b = a_{22}^b = 1$. Then non-cointegration condition Eq. (5) immediately follows. This is another way to express Granger's assertion that at least one way Granger causality is necessary for cointegration. Another implication for the analysis is that under Granger non-causality restrictions, each element of the bivariate non-stationary time series Z_t has a bivariate random walk with drift representation. This is a special case of a more general result, according to which the individual series of a k-dimensional VAR(1) process follow univariate ARMA $(k,k-1)$ models, where k and $k-1$ are the maximum orders of the individual ARMA models.

As mentioned above, Granger causality in at least one direction is necessary for cointegration. However, is it also sufficient? In general, the answer is negative. In particular, for VARs, whose order is greater than one, causality can easily coincide with non-cointegration. However, for first order models and for the majority of the economic time series, causality seems also to be sufficient for cointegration. To show this, assume that $a_{12}^b \neq 0$ and $a_{21}^b = 0$. In such a case, matrix A_1^b is written as

$$A_1^b = \begin{bmatrix} 1 - \alpha_{12}^b \dfrac{\sigma_{xy}(0)}{\sigma_y(0)} & \alpha_{12}^b \\ 0 & 1 \end{bmatrix},$$

where $\sigma_{xy}(0)\cdot t = \mathrm{Cov}(y_t,x_t)$ and $\sigma_y(0)\cdot t = \mathrm{Var}(y_t)$. The trace of A_1^b is:

$$\mathrm{tr}(A_1^b) = 2 - a_{12}^b\frac{\sigma_{yx}(0)}{\sigma_y(0)} = \lambda_1 + \lambda_2 = \lambda_1 + 1. \tag{9}$$

As has been already mentioned, non-cointegration can be thought of as a situation in which both λ_1 and λ_2 being greater than 1. If one of these eigenvalue is less than 1, the system is cointegrated. This, implies, in turn, that under no cointegration the sum of the eigenvalues must be greater than 2, which means that $\lambda_1 \geq 1$ or that $a_{11}^b \geq 1$. This is satisfied when $a_{12}^b\frac{\sigma_{yx}(0)}{\sigma_y(0)} < 0$. On the other hand, the system is cointegrated if $a_{12}^b\frac{\sigma_{yx}(0)}{\sigma_y(0)} > 0$, i.e. when $a_{11}^b < 1$ and thus $\lambda_1 < 1$. The only case in which causality coincides with non-cointegration in the context of a bivariate VAR(1) is when a_{11}^b and/or a_{22}^b are greater than 1. Such cases are not very often encountered in practice. This, in turn, implies that evidence of causality in first-order models usually leads to cointegration. The above conditions clearly indicate that if unit roots were found in the AR representations of y_t and x_t, then they would persist in the VAR representation, provided that there would be no Granger causality between y_t and x_t (*sufficiency*).

To test whether a_{12}^b or a_{21}^b are equal to zero, we follow Johansen's (Johansen, 1988, 1991) approach. By imposing cointegration restrictions, the VAR(1) can be written in the following error correction form:

$$\begin{bmatrix}\Delta y_t\\ \Delta x_t\end{bmatrix} = \begin{bmatrix}\gamma_0\\ \gamma_1\end{bmatrix} + \begin{bmatrix}c_{11}\\ c_{12}\end{bmatrix}\begin{bmatrix}b_{11} & b_{12}\end{bmatrix}\begin{bmatrix}y_{t-1}\\ x_{t-1}\end{bmatrix} + \begin{bmatrix}e_{1t}\\ e_{2t}\end{bmatrix}, \tag{10}$$

where $\mathbf{c}' = [c_{11}\ c_{12}]$ is the 'loading matrix' and $\mathbf{b}' = [b_{11}\ b_{12}]$ is the cointegrating vector. In this framework, the hypothesis $a_{12}^b = 0$ can be regarded as equivalent to the hypothesis $c_{11} = 0$, and the hypothesis $a_{21}^b = 0$ as equivalent to $c_{12} = 0$. Johansen (1992) defines $c_{11} = 0$ as a situation in which Δy_t is weakly exogenous to Eq. (10) when the parameters of interest are the elements of the cointegrating vector. It can be seen that in the context of a bivariate VAR of order one, weak exogeneity is equivalent to the concept of Granger non-causality. Hypotheses on the elements of the loading matrix can be tested by means of likelihood ratio tests, as described in Johansen (1991), once the rank of the matrix $\mathbf{cb}'$ has been determined.

2.1. The trivariate case

In this subsection we extend the above results in the three-variable case and examine the effects of a missing causing variable, say w_t, on the causality structure between y_t and x_t. Consider the following trivariate VAR(1) system:

$$W = A^r + A^r W \quad + E^r, \tag{11}$$

A sufficient condition for Eq. (13) to hold is $a^r_{ij} = 0$ for $i \neq j$, which is equivalent to saying that a necessary condition for cointegration is Granger causality between at least two variables in at least one direction.

As in the bivariate case, non-causality restrictions of the form, $a^r_{ij} = 0$, $i \neq j$, can be tested by means of Johansen's likelihood ratio tests. However, the trivariate case is more complicated than the bivariate one, as far as causality testing in a cointegrated system is concerned. In particular, Eq. (11) can be decomposed by means of

$$\Delta W_t = A^r_0 + \mathbf{c}\mathbf{b}'W_{t-1} + E_t \tag{11b}$$

where $\mathbf{c}$ and $\mathbf{b}$ are $3 \times r$ matrices and r is the dimension of the cointegrating space, i.e. 1 or 2. In such a case, we are interested in testing whether the product of a particular row of matrix c with the corresponding column of matrix $\mathbf{b}'$ is zero. In particular, for $r = 1$ we can define the mapping given below as case 1:

Case 1: $r = 1$

(i) H_0: $c_{11}b_{11} = 0 \Leftrightarrow a^r_{11} = 1$,
(ii) H_0: $c_{11}b_{21} = 0 \Leftrightarrow a^r_{12} = 0$,
(iii) H_0: $c_{11}b_{31} = 0 \Leftrightarrow a^r_{13} = 0$,
(iv) H_0: $\mathrm{c}_{21}b_{11} = 0 \Leftrightarrow a^r_{21} = 0$,
(v) H_0: $c_{21}b_{21} = 0 \Leftrightarrow a^r_{22} = 1$,
(vi) H_0: $c_{21}b_{31} = 0 \Leftrightarrow a^r_{23} = 0$,
(vii) H_0: $c_{31}b_{11} = 0 \Leftrightarrow a^r_{31} = 0$,
(viii) H_0: $c_{31}b_{21} = 0 \Leftrightarrow a^r_{32} = 0$,
(ix) H_0: $c_{31}b_{31} = 0 \Leftrightarrow a^r_{33} = 1$.

For $r \geq 2$ and $r = 0$ we distinguish among the following cases:

Case 2: $r = 2$.

Recently, Hunter (1990) and Mosconi and Giannini (1992) consider another type of restriction, particularly interesting in the context of first-order models with cointegration rank equal to two. For instance, this type of restriction can take the form $c_{11}b_{21} + c_{12}b_{22} = 0$, which in the context of the trivariate model with $r = 2$, is equivalent to testing for $a^r_{12} = 0$. In general, this type of restriction allows for testing individual non-causality restrictions, $a^r_{ij} = 0$, in the context of a cointegrated VAR, with $r(\Pi) = 2$, $\Pi = \mathbf{cb}'$.

Case 3: $r = 3$.

In such a case we have a stable VAR in levels and standard F-tests can be used for the hypothesis $a^r_{ij} = 0$.

Case 4: $r = 0$.

In such a case we have an unstable, non-cointegrated VAR, where no Granger causality is expected to be found, unless a^r_{11}, a^r_{22}, or a^r_{33}, are greater than one. Again, standard F-tests in a first differenced VAR can be used to confirm that ΔW_t in Eq. (11) is a vector white noise.

At this point the following question may arise: Why should we prefer to test for non-causality restrictions in the context of a cointegrated vector error correction model instead of a VAR in levels? The issue of testing for causality in a non-stationary, possibly cointegrated, VAR was initially addressed by Sims et al. (1990) in the context of a trivariate VAR and then examined in a more general setting by Toda and Phillips (1994). The results from these studies can be summarized as follows: Wald test statistics for non-causality restrictions in the context of the unrestricted VAR will have, in general, non-standard limiting distributions in which nuisance parameters are also present. The Wald test will have a χ^2 asymptotic distribution, free of nuisance parameters, only if there is 'sufficient' cointegration with respect to the variables whose causal effects are being tested (Toda and Phillips, 1994). In such a case, the coefficients of interest appear as coefficients on zero mean stationary regressors in a regression that includes a constant, which is necessary and sufficient for the Wald test to have a limiting χ^2 distribution (Sims et al., 1990). This in turn implies that in order to decide which asymptotic theory applies to a particular problem, some prior information on the presence and location of the unit roots in the VAR is required. This information, however, is difficult to obtain from the estimation of a levels VAR. The picture, however, seems more promising when causality testing is being conducted in the context of the ECM formulation Eq. (11b), once prior information on the cointegration rank has been utilized. This is because the conditions that ensure standard χ^2 asymptotics are usually met. For example, in the case of a trivariate VAR with a single cointegrating vector, a sufficient condition for standard mixed normal asymptotics amounts to the causing variable entering the unique cointegrating vector (see Toda and Phillips, 1994).

Finally, some comments on the restrictiveness of the above analytical framework are in order. First, the Markov framework that results in first-order models is mainly employed in the empirical literature on unit roots and cointegration, where one unit root is widely used. Nevertheless, the 'one-stage' dependence assumption may be quite restrictive, especially in applications using quarterly or monthly data. Therefore this assumption has to be tested before any valid inference can be drawn. Even in the cases in which there is no evidence for unmodeled higher order dynamics, it is advisable to check the sensitivity of any causality — cointegration findings prior to the selection of the order of the system. Second, in the context of alternative systems, the analytical benefit from the adoption of a Markov framework is that we can investigate in greater depth the linkages between unit roots, cointegration and causality. The main results still remain valid. Nonetheless, the most important change concerns one way Granger causality, which is only necessary for cointegration even in non-explosive situations. That is, in the absence of cointegration, ΔW_t is not a vector white noise. In such a situation the direction of causality can be decided upon via standard F-tests in the first differenced VAR.

2.2. Causality inference in incomplete systems

Finally, it is worth mentioning the sensitivity of the causality inference between y_t and x_t in the case of an omitted variable w_t. It has been shown elsewhere (Caporale and Pittis, 1997) that the 'role' of w_t in the extended system (whether it causes y_t and/or x_t or is caused by them) is the sole determining factor for the sensitivity of causality inference between y_t and x_t. Let us focus on this issue, by adopting alternative scenaria about the role of the 'omitted' variable w_t. As already mentioned, in the following notation superscripts 'r' and 'b' refer to the trivariate and the bivariate system, respectively.

Case I: w_t is caused by both y_t and x_t, i.e. $\alpha_{31}^r, \alpha_{32}^r \neq 0$.

(a) Suppose that w_t causes both y_t and x_t, i.e. $\alpha_{13}^r, \alpha_{23}^r \neq 0$: In this case, by comparing Eqs. (3a)–(3d) and Eqs. (12a)–(12i), we immediately see that $\alpha_{12}^b \neq \alpha_{12}^r$ and $\alpha_{21}^b \neq \alpha_{21}^r$. Therefore in the case that the 'omitted' variable has a causing effect on both variables of the incomplete system, the causality inference in both directions is wrong.
(b) w_t causes x_t but not y_t, i.e. $\alpha_{13}^r = 0$, $\alpha_{23}^r \neq 0$. Here we get $\alpha_{12}^b = \alpha_{12}^r$ and $\alpha_{21}^b = \alpha_{21}^r$. Therefore in the case that the 'omitted' variable has a causing effect on only one of the variables of the incomplete system, i.e. on x_t, the causality inference in one direction remains valid ($x_t \rightarrow y_t$) but it becomes invalid in the opposite direction $y_t \rightarrow x_t$).
(c) w_t causes neither y_t nor x_t. In this case $\alpha_{12}^b = \alpha_{12}^r$ and $\alpha_{21}^b = \alpha_{21}^r$, which means that if the omitted variable does not cause any of the variables of the incomplete system, then, within that system, the causality inference in both directions remains valid. It must be noted that, even in this case, the bivariate system remains an 'incomplete' system, since we have allowed for w_t to be caused by both y_t and x_t.

Case II: w_t is not caused by either y_t or x_t, i.e. $\alpha_{31}^r =_{32}^r = 0$.

Again, by comparing Eqs. (3a)–(3d) and Eqs. (12a)–(12i), we can see that the results concerning the causality inference between (y_t, x_t), within the incomplete system, are similar to those associated with the cases (a)–(c) described above. We should only note that in the case (IIc), the bivariate system should not be referred to as an incomplete system since the subvector $[y_t, x_t]'$ is independent of w_t. Therefore we can conclude that cases (a)–(c) hold regardless of whether w_t is caused by y_t and x_t. Finally, case (c) describes a necessary and sufficient condition for the causality inference between y_t and x_t to be invariant to the system within which it is examined.

3. Empirical analysis

The preceding theoretical analysis seems to suggest the following empirical approach:

(1) Carry out univariate tests for unit roots to check whether each of the series y_t and x_t contains a unit root. If this is the case, proceed to step 2:

(2) Estimate a bivariate VAR(1) in levels and test whether the Markovness assumption is valid for the data in hand. This can be carried out by testing for the statistical adequacy of this model and by comparing it with higher-order, competing in terms of their information criteria, models (Akaike, 1973, Schwarz, 1978). If the

estimates of α_{11} and α_{12}, i.e. $\hat{\alpha}_{11}$ and $\hat{\alpha}_{22}$, are close to one, we should expect no evidence of Granger causality in any direction ($\alpha_{12} = \alpha_{21} = 0$);[2] this is a sufficient condition for no cointegration between y_t and x_t. If, on the other hand, either $\hat{\alpha}_{11}$ or $\hat{\alpha}_{22}$ are much lower than one, then at least one way Granger causality should be present, which in turn implies that the necessity (but, in general, not the sufficiency), for cointegration is fulfilled. Given the nature of the time-series employed here, it is sensible to argue that causality will almost surely lead to cointegration, since explosive situations of the type α_{11}, $\alpha_{22} > 1$ are highly unlikely to occur.

(3) Next, perform formal cointegration tests between y_t and x_t. Here distinguish between two cases: (a) *Cointegration*: This case formally establishes the presence of Granger causality in at least one direction and as such it should coincide with an estimate of α_{11} or $\alpha_{22} < 1$. The direction of causality should be decided upon testing $\alpha_{12} = 0$ or $\alpha_{21} = 0$ by means of Johansen's likelihood ratio tests on the elements of the loading matrix **c**, as described in the previous section. (b) *Non-Cointegration*: In general, this case coincides with either no causality ($\alpha_{11},\alpha_{22} \cong 1$) or with causality (at least one of the estimates of α_{11} or $\alpha_{22} > 1$) since both situations satisfy Eq. (5). However, evidence of non-cointegration, in the present context, must imply the absence of causality in any direction since European interest rates have not exhibited explosive behavior. In such a case and under the maintained hypothesis of Markovness, both Δy_t and Δx_t are white noises. This hypothesis can be tested by means of standard F-tests in the context of a first-differenced higher-order VAR.

(4) The procedure discussed below should be undertaken if the bivariate analysis provides evidence against cointegration. Recent developments in testing for unit roots in a univariate or a multivariate framework, have shown that a once and for all exogenous shock to the deterministic component (mean or trend) of a particular series can bias the standard unit root tests towards non-rejection of the null, with the degree of the bias being rising with the magnitude of the shock (Perron, 1989, 1990; Banerjee and Urga, 1995). Therefore before looking for other reasons that might have led to non-cointegration in a bivariate system, we should examine whether such deterministic breaks are responsible for the non-rejection of the null. Potential break dates that are also economically interpretable include the realignment dates of each individual EMS currency and the August 1993 episode, during which the EMS fluctuation bands were widened. Sequential tests, for identifying these potential break dates are also available. The recursive sum of squared residuals is very useful in revealing such dates. A more formal testing procedure is suggested (Banerjee and Urga, 1995), inter alia in the form of the sequence of one-step ahead Chow tests at each period i for the hypothesis that no structural break has occurred within the sample of size i (see Appendix A). If the economically interpretable break dates coincide with those suggested by the tests, then the use of dummy variables is well justified. Hence, cointegration is reexamined by

[2] For notational simplicity, the superscripts b and r have been dropped, since there is no risk of confusion.

estimating the bivariate system with dummy variables, taking the value one at the break dates and zero elsewhere. If non-cointegration still characterizes the bivariate system, then proceed to step 5.

(5) The analysis of causality in incomplete systems (Section 2.2) has shown that the omission of a ‘causing’ variable can affect causality (and thus cointegration) inference between the variables y_t and x_t of a bivariate system. Therefore before reaching any conclusion on the linkages between y_t and x_t, causality and cointegration analysis, as described in the previous steps, should be repeated in a trivariate context.

(6) As already mentioned, the preceding analysis is carried out in a Markov environment. Even if there is no evidence for unmodeled higher order dynamics from steps 2 and 3, it would be advisable to check for the robustness of the cointegration findings to the selection of alternative orders (especially whether the rank of $\mathbf{cb}'$ remains unchanged).

(7) Hypotheses testing, in a bivariate and a trivariate context, apropos of the asymmetric functioning of the EMS are tabulated below. They concern alternative versions of the GDH and other hypotheses in relation to the determination of the short-term interest rate. As far as notational correspondence with the theoretical section is concerned, y_t represents the interest rate of each individual European Union country, and x_t and w_t are the German and the US interest rates, respectively. Similarly, subscripts 1, 2 and 3 stand for Germany, other European Union countries taken one at a time, and USA, respectively. The hypotheses at issue are:

I: Bivariate Context

- (i) German Dominance Hypothesis (GDH): $\alpha_{12} \neq 0$ and $\alpha_{21} = 0$,
- (ii) Symmetry: $\alpha_{12} \neq 0$ and $\alpha_{21} \neq 0$,
- (iii) Monetary autonomy in both countries: $\alpha_{12} = 0$ and $\alpha_{21} = 0$.

II: Trivariate Context

- (iv) Strong GDH (*no direct or indirect causality from US*): $\alpha_{12} \neq 0$, $\alpha_{13} = 0$, $\alpha_{23} = 0$,
- (v) Weak GDH of type 1 (*direct causality from US*): $\alpha_{12} \neq 0$, $\alpha_{13} \neq 0$, $\alpha_{23} = 0$,
- (vi) Weak GDH of type 2 (*direct and indirect causality from US*): $\alpha_{12} \neq 0$, $\alpha_{13} \neq 0$, $\alpha_{23} \neq 0$,
- (vii) Semi strong GDH (*no direct causality from US; only indirect causality through Germany*): $\alpha_{12} \neq 0$, $\alpha_{13} = 0$, $\alpha_{23} \neq 0$,
- (viii) US Dominance Hypothesis (*direct causality from the US with no direct causality from Germany*): $\alpha_{12} = 0$, $\alpha_{13} \neq 0$, $\alpha_{23} \neq 0$.

As already mentioned, all the above hypotheses have to be translated in terms of the products of the elements c_{ij} of the matrix of the loading factors, with the elements of the cointegrating vectors, as described in Section 2.1. The data used in the analysis are quarterly short-term (on-shore) market interest rates taken from OECD’s *Main Economic Indicators*. They concern Belgium, France, Germany, Ireland, Italy, the Netherlands and the United States, and cover the period from 1979.3 to 1994.4.

3.1. Unit roots in the AR(1) representations of the series

The first step in examining the statistical properties of the time series employed here is to test for the presence of unit roots in the autoregressive representation of the series in hand. Table 1 reports the estimates from Dickey-Fuller (DF) and augmented Dickey-Fuller (ADF) tests of the null hypothesis that a unit root exists. Both tests fail to reject the unit-root null, even when dummy variables are included to account for once-off exogenous changes, such as realignments to the deterministic component of the series (Perron, 1989). The point estimates of the AR(1) coefficient are also reported in this table (column 1) to facilitate a direct comparison of these estimates with the estimates concerning the diagonal elements of matrix A in our VAR models of Section 2.1 (Table 2).

3.2. Bivariate analysis

Next, we estimate bivariate VAR(1) [BVAR(1)] models consisting of the interest rate of each individual EMS country and that of Germany. Estimation results along with suitable misspecification tests are reported in Table 2. The choice of first-order models is well supported both by the misspecification tests, which reveal no departures from the model assumptions, and by the Schwarz information criterion. The latter achieves its lowest value for a lag length equal to one. The estimates of α_{11} and α_{22} are close to one for all countries except for the Netherlands. That is, according to the analysis of Section 2, in all bivariate systems at issue, except that for the Netherlands, Granger causality in any direction is expected to be absent. This implies that these systems are not expected to be cointegrated. The case of the Netherlands is different. The respective estimates of α_{11} and $\alpha_{22} = 0.59$ and 0.90, respectively. Therefore we should expect causality to run from Germany to

Table 1
Estimates of the AR(1) coefficients and unit root tests

Countries	AR(1) coeff.	DF	ADF(1)	ADF(2)	ADF(3)	ADF(4)
BE	0.941 (0.051)	−1.66	−1.84	−1.56	−1.20	−1.31
FR	0.920 (0.054)	−1.03	−1.30	−1.05	−1.02	−1.14
GE	0.933 (0.043)	−1.31	−2.08	−1.93	−1.91	−2.32
IR	0.873 (0.069)	−2.20	−2.39	−1.78	−1.79	−1.64
IT	0.945 (0.042)	−0.78	−0.86	−0.83	−0.89	−0.91
NL	0.909 (0.059)	−1.64	−1.95	−1.83	−2.01	−1.93
US	0.921 (0.063)	−1.24	−1.31	−0.99	−1.27	−0.97

Notes. (1) Critical values for the DF and ADF statistics are: 5% = −2.914; 1% = −3.55. Numbers in parentheses are standard errors. (2) The Country codes are: BE, Belgium; FR, France; GE, Germany; IR, Ireland; IT, Italy; NL, Netherlands; and US, United States.

Table 2
Bivariate analysis: VAR(1) in levels

	BE	GE	FR	GE	IR	GE	IT	GE	NL	GE
α_{11}	0.921		0.835		0.773		0.969		0.598	
	(0.061)		(0.051)		(0.071)		(0.032)		(0.112)	
α_{12}	0.052		0.061		0.229		0.068		0.327	
	(0.061)		(0.060)		(0.091)		(0.051)		(0.102)	
α_{21}		−0.028		−0.112		−0.101		−0.065		0.042
		(0.052)		(0.052)		(0.042)		(0.032)		(0.113)
α_{22}		0.931		0.978		1.000		0.958		0.901
		(0.057)		(0.056)		(0.053)		(0.054)		(0.102)
P-values for misspecification tests										
AR(4)	0.615	0.912	0.051	0.104	0.285	0.011*	0.225	0.085	0.345	0.092
NORM	0.587	0.059	0.01*	0.102	0.033*	0.312	0.01*	0.074	0.063	0.135
ARCH	0.923	0.726	0.897	0.099	0.021*	0.182	0.313	0.688	0.452	0.087
LIN	0.165	0.558	0.396	0.978	0.405	0.489	0.186	0.783	0.128	0.567
SIC (4)	−0.73		−0.70		0.80		0.26		−1.60	
SIC (3)	−0.77		−0.93		0.57		0.05		−1.77	
SIC (2)	−0.93		−1.16		0.31		−0.14		−1.79	
SIC (1)	−0.92		−1.33		0.11		−0.21		−2.01	

Notes. (1) A description of the misspecification tests is given in the Appendix. (2) An asterisk denotes rejection of the null at the 5% significance level. (3) SIC(i) stands for the Schwarz Information Criterion for lag i.

the Netherlands ($\alpha_{12} \neq 0$). This is a necessary condition for cointegration. Furthermore, as shown in Section 2, for non-explosive series in the context of BVAR(1) models, Granger causality in one direction is also sufficient for cointegration. In view of the above, we proceed to carry out formal cointegration tests.

Johansen's cointegration tests are reported in Table 3. These results confirm the preceding finding that the Netherlands is the only EMS country cointegrated with Germany. They are also consistent with the findings of earlier studies (Karfakis and Moschos, 1990, Katsimbris and Miller, 1993). These studies utilize monthly data (1979.4–1988.11) and report negative evidence on bilateral cointegration for all EMS countries except for the Netherlands. Concerning the latter, cointegration implies the existence of Granger causality in at least one direction. To identify the direction of causality, we have performed likelihood ratio tests on the elements of the matrix **c**, as described in Section 2. The hypothesis that the German interest rate does not Granger cause the Dutch rate is soundly rejected [$\chi^2(1) = 11.69$], whereas the hypothesis that no Granger causality runs from the opposite direction is easily accepted [$\chi^2(1) = 0.26$]. This finding establishes the validity of a strong German Dominance Hypothesis in the case of the Netherlands.

The next question is whether the non-cointegration finding of the other bivariate systems under consideration has remained unaltered over the whole sample period. In fact, the EMS has gone through different phases since its creation in 1979. A period of large and frequent realignments, during the early eighties, followed by a period of relative tranquillity, was succeeded by a turbulent period leading to a widening of the fluctuation bands from $\pm 2.25\%$ since its inception to $\pm 15\%$ in August 1993. It is, therefore, proper to question the time-invariance of the causality and cointegration features of the interest rates under consideration. To investigate the issue, we have estimated the five BVAR(1) models recursively, and examined the time profile of the maximum eigenvalue of the matrix Π for each

Table 3
Bivariate cointegration tests between the German and the EMS interest rates

Country	Max eigenvalue		λ-max		Trace	
	D	ND	D	ND	D	ND
BE	0.08	0.05	5.16	3.57	6.91	5.74
FR	0.18	0.11	12.49	7.21	13.87	11.77
IR	0.19	0.17	11.67	10.35	15.27	14.85
IT	0.13	0.12	8.11	7.87	14.26	9.85
NL	0.38*	0.31*	28.08*	24.69*	30.23*	29.56*

Notes. (1) Trace and λ-max stand for the trace and the maximum eigenvalue statistics, respectively (Johansen, 1988). (2) An asterisk indicates rejection of the null hypothesis of no-cointegration at the 5% level. Critical values for λ-max are 14.1 and for the trace 15.4 at the 5% level (Osterwald-Lenum, 1992). (3) D and ND stand for dummies and no-dummies, respectively. The dummies used for each system take the value one for the following periods: BE: October 81, February 82, June 82, August 93; FR: October 81, June 82, March 83, April 86, August 93; IR: October 81, June 82, March 83, August 93; IT: October 81, June 82, July 85, September 92. Otherwise, they take the value zero.

bivariate system. Such estimates are reported in Fig. 1a–e as curves I. There it can be seen that these estimates have remained roughly stable at low levels (in most cases around 0.2) till August 1993, when they declined further. This means that the non-cointegration property that characterizes all the BVAR(1) models, except that of the Netherlands, is time invariant. Moreover, the 'cointegrability' of these systems has decreased after the widening of the fluctuation bands, indicating the laxity of linkages of short-run interest rates within the EMS. A notable exception is, again, the Netherlands: the respective maximum eigenvalue estimates have not declined. This suggests that the Dutch monetary authorities, did not choose to increase the degree of their monetary autonomy, despite the fact that the new institutional arrangement had allowed them to do so. The case of Italy is slightly different as well. The Italian interest rate seems to have been cointegrated with the German one, up to approximately mid-1982, when the bilateral exchange rate for the Deutsche mark against the Italian lira rose by 7%. This finding reinforces our point [Section 2, methodological step Eq. (4)], that any failure to reject the null hypothesis of no cointegration should be interpreted with care, if it is known that distinct once-and-for-all breaks have occurred within the sample period under consideration. That is, before one accepts without any reservation the non-cointegration results, he must ensure that the non-rejection of the null is not due to deterministic breaks in his series. Sequential tests for identifying potential break dates have been carried out and are described in the appendix. These tests can enable us to reject the null hypothesis of structural invariance, exactly at the economically interpretable dates of realignments. To that end, we have repeated the cointegration tests with dummy variables included in the BVAR(1) models. The respective results are also reported in Table 3, with the null hypothesis of no cointegration still surviving the evidence. However, it can be seen that the inclusion of dummy variables, for the relevant realignment dates, has increased throughout the estimates of the maximum eigenvalue of the matrix Π. This can be seen in Fig. 1a–e, where the recursive estimates of the maximum eigenvalues, in the presence of dummies, are reported as curves II. Hence, the realignments have contributed to the non-cointegrability of the systems in question, although they are not solely responsible for that.

The overall conclusion emerging from our bivariate analysis is that the ERM does not seem to have induced a link between the interest rates of the EMS countries and that of Germany. The only exception appears to be the Netherlands for which non-cointegration is easily rejected. This finding is hardly surprising. For in the history of the EMS there has not been a single realignment in which the Dutch guilder has not revalued pari passu with the Deutsche mark against the other EMS currencies.

On the other hand, non-cointegration between the rest of the EMS rates and the German one, could merely be the result of gradual convergence of the interest rates of individual EMS countries towards the German rate (see Caporale and Pittis, 1993). This in turn implies that no comovement or long run relationship should be expected to prevail in a sample dominated by a 'convergence period'. This is a point which most of the relevant studies fail to recognize.

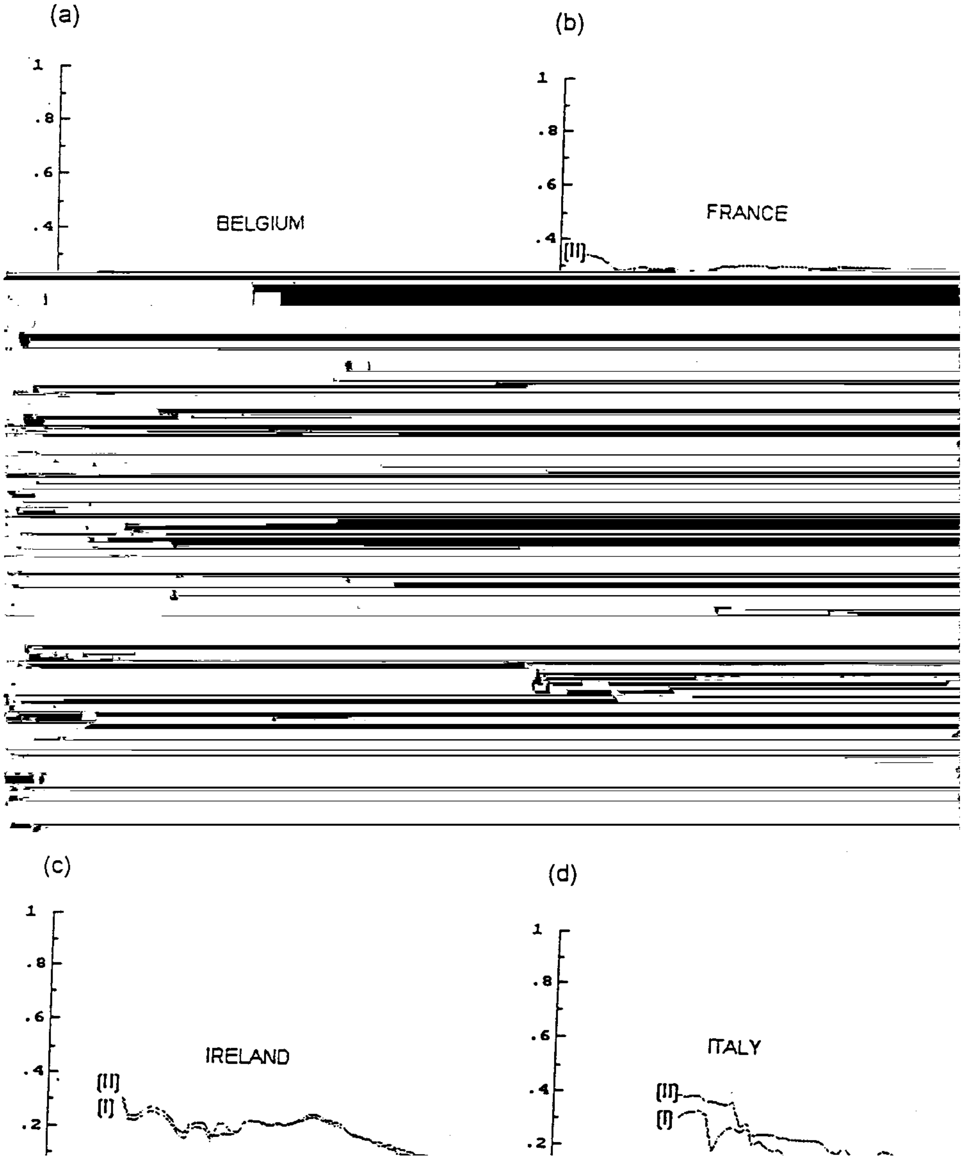

Fig. 1. Recursive estimates of the maximum eigenvalue in the bivariate systems.

3.3. Trivariate analysis

Our analysis on causality in incomplete systems (Section 2) has demonstrated that causality inference and, hence, cointegration is strongly affected by the omission of a third variable, which causes one or both of the existing variables. As already mentioned, the US short-term interest rate seems to be an obvious candidate for such a variable, since the US monetary policy has likely affected the policies of a number of countries, including the EMS countries.

Estimation results for the trivariate VAR(1) models [TVAR(1)] along with their misspecification tests are reported in Table 4. Again, no departure from the underlying model assumptions is detected. Moreover, the Schwartz information criterion seems to suggest that the first order models, for all countries, are preferable to the higher-order ones. Interestingly enough the estimates of the α_{11} coefficient appear to be significantly lower than one for all the countries under consideration. This implies that Granger causality is expected to run from either Germany ($\alpha_{12} \neq 0$) or the US ($\alpha_{13} \neq 0$) or both to each individual EMS country. In addition, in all cases, the estimates of α_{22} also appear to be smaller than unity, although they are greater in size than the respective α_{11} estimates. This suggests that either each EMS interest rate or the US rate or both should cause the German rate [see Eqs. (12a)–(12i)]. Finally, in all cases, the α_{33} estimates are indistinguishably different from one. This result is pointing towards no Granger causality running from any EMS country and/or Germany to the US.

Since Granger causality has been informally detected to be present in all trivariate systems, the latter are expected to be cointegrated. Formal cointegration tests are given in Table 5. They reveal that each EMS country's interest rate is cointegrated with the German and the US interest rate in a trivariate framework, with the dimension of the cointegration space being equal to one. This however, does not necessarily imply that all three variables in each system enter the cointegrating vector $\mathbf{b} = [b_{11}, b_{21}, b_{31}]'$. For example, assume that the long-run relationship among the interest rates in question, is given by equation:

$$b_{11}i_t^* + b_{21}i_t^{\mathrm{GE}} + b_{31}i_t^{\mathrm{US}} = 0 \tag{14}$$

There is no a priori guarantee that all elements of $\mathbf{b}$ are different from zero. For example, assume that $b_{21} = 0$. In such a case, the one cointegrating vector, detected by the Johansen tests, reflects a long-run relationship between the interest rate of the particular EMS country (i_t^*) and the US rate (the German rate does not enter the relationship). Testing for such hypotheses are of paramount importance in the context of the questions raised in this paper and can be carried out by means of likelihood ratio tests (Johansen, 1991). Table 6a reports the results from testing for zero restrictions on the elements of both the cointegrating vector $\mathbf{b}$ and the loading matrix $\mathbf{c} = [c_{11}, c_{21}, c_{31}]'$ for the whole sample. A rather provocative result emerges from this exercise: the German interest rate does not enter the cointegrating relationship for any EMS country except Belgium. This automatically precludes the possibility that the previously detected Granger causality is of the type $\alpha_{12} \neq 0$, which is evidence against any type of the GDH. Moreover, the hypothesis $\alpha_{31} = 0$ cannot be rejected for any of the systems under consideration. This in turn implies that the US interest rate is weakly exogenous in all systems, when the elements of the cointegrating vector are considered (Johansen, 1992). The results in Table 6a suggest that matrix A_1^r, which describes the causality

Table 4
Trivariate analysis: VAR(1) in levels

	BE	GE	US	FR	GE	US	IR	GE	US	IT	GE	US
α_{11}	0.541			0.793			0.711			0.798		
	(0.083)			(0.064)			(0.065)			(0.045)		
α_{12}	0.225			0.102			0.146			0.055		
	(0.062)			(0.051)			(0.076)			(0.041)		
α_{13}	0.251			0.061			0.246			0.246		
	(0.052)			(0.042)			(0.067)			(0.048)		
α_{21}		−0.312			−0.275			−0.187			−0.135	
		(0.063)			(0.056)			(0.046)			(0.035)	
α_{22}		0.812			0.892			0.872			0.857	
		(0.041)			(0.041)			(0.054)			(0.048)	
α_{23}		0.198			0.169			0.146			0.162	
		(0.031)			(0.038)			(0.042)			(0.033)	
α_{31}			0.038			−0.018			−0.089			0.027
			(0.125)			(0.088)			(0.055)			(0.066)
α_{32}			−0.021			−0.091			−0.072			−0.012
			(0.092)			(0.063)			(0.062)			(0.074)
α_{33}			1.000			0.978			1.000			1.000
			(0.075)			(0.055)			(0.055)			(0.068)
P-values for misspecification tests												
AR(4)	0.566	0.473	0.756	0.634	0.472	0.653	0.356	0.398	0.225	0.448	0.755	0.165
NORM	0.334	0.706	0.053	0.875	0.703	0.068	0.387	0.102	0.043*	0.178	0.968	0.187
ARCH	0.915	0.914	0.985	0.911	0.854	0.452	0.088	0.941	0.812	0.199	0.099	0.179
LIN	0.456	0.798	0.712	0.698	0.952	0.607	0.275	0.280	0.240	0.456	0.198	0.616
SIC (4)	1.26			0.39			2.13			1.19		
SIC (3)	0.71			−0.21			1.63			0.78		
SIC (2)	0.41			−0.03			1.39			0.62		
SIC (1)	−0.22			−0.51			0.93			0.09		

Notes. (1) A description of the misspecification tests is given in the Appendix. (2) An asterisk denotes rejection of the null at the 5% significance level. (3) SIC(i) stands for the Schwarz Information Criterion for lag i.

Table 5
Trivariate cointegration tests

Countries	Max eigenvalue	λ-max	Trace
BE	0.54	48.11*	51.63*
FR	0.37	28.88*	39.23*
IR	0.55	42.35*	51.71*
IT	0.59	53.55*	56.85*

Notes. Critical values for λ-max and for trace are 21.0 and 29.7, respectively (Osterwald-Lenum, 1992).

linkages of interest, takes a different form for Belgium on the one hand, and a different one for France, Ireland and Italy on the other. In particular, we write:

Matrix A_1^r for Belgium; Matrix A_1^r for France, Ireland, Italy

$$\begin{bmatrix} \alpha_{11} & \alpha_{12} & \alpha_{13} \\ \alpha_{21} & \alpha_{22} & \alpha_{23} \\ 0 & 0 & 1 \end{bmatrix}, \qquad \begin{bmatrix} \alpha_{11} & 0 & \alpha_{13} \\ \alpha_{21} & 1 & \alpha_{23} \\ 0 & 0 & 1 \end{bmatrix}.$$

In all cases, the US interest rate seems to cause all the European rates but is not caused by them. This finding contradicts the findings by Artus et al. (1991) and Kirchgaessner and Wolters (1993). Concerning the intra-EMS linkages, the case of Belgium is more ‘normal’; it reflects a symmetric functioning of the monetary policies of Belgium and Germany. The other three cases are rather ‘strange’: The German rate appears to be caused by rather than cause the interest rates of France, Ireland or Italy. This result may be counterintuitive at a first glance. However, the concurrence of (a) the escalation of the German interest rates due to Bundesbank’s tight monetary policy following the German reunification; (b) the increasing uncertainty as regards the future of the EMS (the British pound and the Italian lira withdrew from the ERM in September 1992); and (c) the resulting widening of the EMS fluctuation bands in August 1993 seem to have induced, as already mentioned, a higher degree of monetary autonomy to the system. In addition, the preceding period was characterized by a high degree of uncertainty as regards the future of the system, since both the British pound and the Italian lira withdrew from the ERM in September 1992. That turbulent period might have affected the causality linkages of the interest rates under consideration. Indeed, the dynamic properties of all trivariate systems have been significantly affected since the beginning of 1991. This can be seen in Fig. 2a–d, where recursive estimates for the maximum eigenvalue of the systems under consideration are presented. The evidence indicates that the long-run properties of the systems seem to have been considerably stable during the 1987–1991 period. Nonetheless, a distinct drop of the maximum eigenvalue has occurred in the first quarter of 1991. Since that date the maximum eigenvalue for all systems has exhibited an erratic behavior. It is natural then to try to investigate the causality linkages among the

Table 6
Testing for zero restrictions on the cointegrating and loading vectors

		BE	FR	IR	IT
(a): *Sample 79.2–94.4*					
H_0: $c_{11} = 0$	$\chi^2(1)$	25.771	7.175	20.392	29.135
	P value	(0.000)*	(0.007)*	(0.000)*	(0.000)*
H_0: $c_{21} = 0$	$\chi^2(1)$	31.075	20.712	17.865	24.748
	P value	(0.000)*	(0.000)*	(0.000)*	(0.000)*
H_0: $c_{31} = 0$	$\chi^2(1)$	0.012	0.471	2.745	0.005
	P value	(0.981)	(0.495)	(0.104)	(0.935)
H_0: $b_{11} = 0$	$\chi^2(1)$	40.051	18.567	33.143	35.789
	P value	(0.000)*	(0.000)*	(0.000)*	(0.000)*
H_0: $b_{21} = 0$	$\chi^2(1)$	13.210	1.553	1.182	0.028
	P value	(0.000)*	(0.213)	(0.275)	(0.861)
H_0: $b_{31} = 0$	$\chi^2(1)$	41.415	16.897	25.856	49.951
	P value	(0.000)*	(0.000)*	(0.000)*	(0.000)*
(b): *Sample 79.2–91.1*					
H_0: $c_{11} = 0$	$\chi^2(1)$	24.287	12.276	16.522	25.258
	P value	(0.000)*	(0.000)*	(0.000)*	(0.000)*
H_0: $c_{21} = 0$	$\chi^2(1)$	19.465	19.921	10.432	14.672
	P value	(0.000)*	(0.000)*	(0.000)*	(0.000)*
H_0: $c_{31} = 0$	$\chi^2(1)$	0.319	0.199	0.215	0.653
	P value	(0.572)	(0.655)	(0.604)	(0.418)
H_0: $b_{11} = 0$	$\chi^2(1)$	37.742	25.113	28.654	32.711
	P value	(0.000)*	(0.000)*	(0.000)*	(0.000)*
H_0: $b_{21} = 0$	$\chi^2(1)$	14.815	5.837	4.659	6.734
	P value	(0.000)*	(0.015)*	(0.022)*	(0.012)*
H_0: $b_{31} = 0$	$\chi^2(1)$	32.637	11.134	18.568	38.313
	P value	(0.000)*	(0.000)*	(0.000)*	(0.000)*
(c): *Sample 79.2–88.4*					
H_0: $c_{11} = 0$	$\chi^2(1)$	13.598	7.283	9.313	4.876
	P value	(0.000)*	(0.007)*	(0.002)*	(0.027)*
H_0: $c_{21} = 0$	$\chi^2(1)$	19.831	37.92	15.059	16.472
	P value	(0.000)*	(0.000)*	(0.000)*	(0.000)*
H_0: $c_{31} = 0$	$\chi^2(1)$	1.649	0.058	0.511	2.687
	P value	(0.199)	(0.808)	(0.474)	(0.101)
H_0: $b_{11} = 0$	$\chi^2(1)$	24.446	14.943	11.345	11.607
	P value	(0.000)*	(0.000)*	(0.000)*	(0.000)*
H_0: $b_{21} = 0$	$\chi^2(1)$	10.335	15.581	9.339	12.519
	P value	(0.001)*	(0.000)*	(0.002)*	(0.000)*
H_0: $b_{31} = 0$	$\chi^2(1)$	56.849	34.639	38.994	28.664
	P value	(0.000)*	(0.000)*	(0.000)*	(0.000)*

interest rates under consideration, for the time span 1979.2–1991.1. The results from testing for zero restrictions in both the cointegrating vector and the vector of the loading factors are reported in Table 6b. The evidence suggests that the German rate enters into the cointegrating vector of all systems under consideration. In addition, the hypothesis that the US rate is weakly exogenous with respect

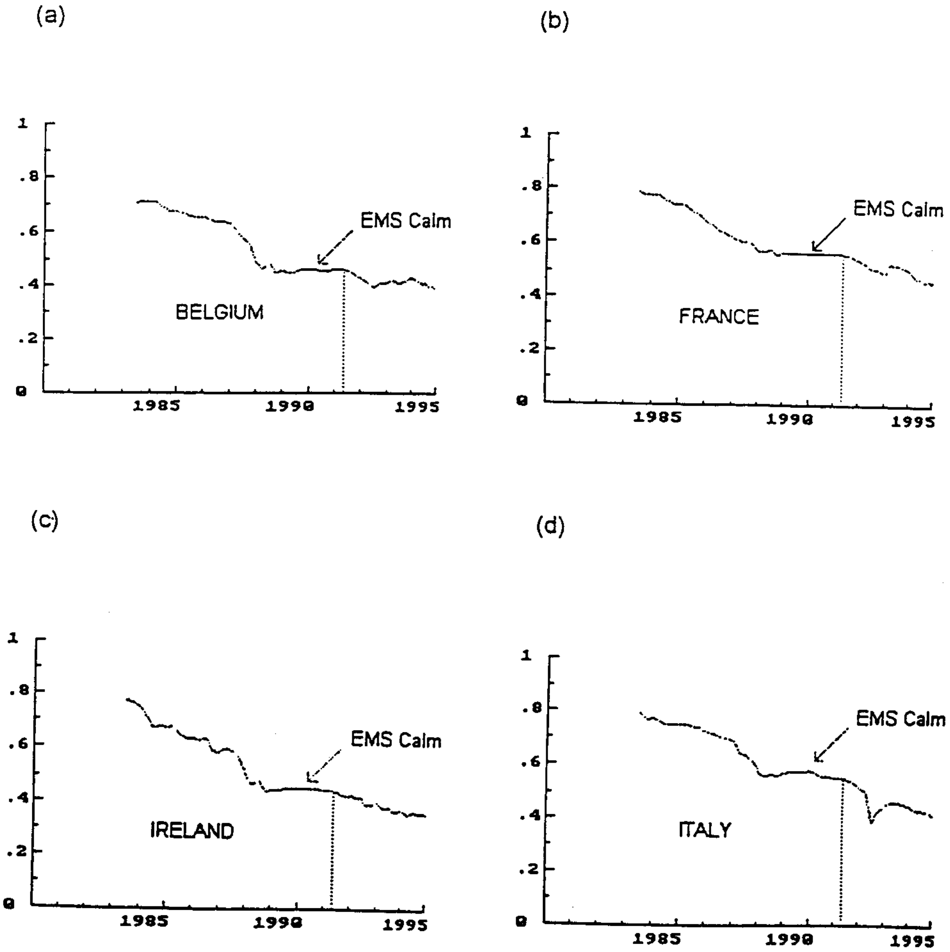

Fig. 2. Recursive estimates of the maximum eigenvalue in the trivariate systems.

to the cointegrating vector ($c_{31} = 0$) cannot be rejected. Hence, matrix A_1^r takes the same form for all systems, i.e.

$$A_1^r = \begin{bmatrix} \alpha_{11} & \alpha_{12} & \alpha_{13} \\ \alpha_{21} & \alpha_{22} & \alpha_{23} \\ 0 & 0 & 1 \end{bmatrix}.$$

A general result emerging from the empirical analysis concerns the performance of the US interest rate in Europe. In particular, it appears that, during our sample period, the US rate has caused all European rates, including the German one, but has not been caused by anyone of them. This result is in agreement with the common arbitrage condition on capital markets, according to which the ex ante real interest rates should be equal across the world. In such a situation two shocks can occur: At first, monetary shocks: Suppose monetary policy becomes tighter in the US. Real US interest rates rise. In order for arbitrage opportunities not to arise, real European interest rates must also rise. Secondly, suppose inflation expectations rise in the US. These are translated into higher nominal US interest rates, without necessarily affecting the real rates. The higher inflation expectations in the US can be transmitted into Europe via trade relations. The European imports from the US become more expensive, thus resulting in imported inflation. If trade relations are intensive and European import demand for American exports is income inelastic, these can induce interest rates in Europe to rise as well.

Concerning now the question of a unidirectional causality from the US to Europe, we can argue as follows: So far, the US, an essentially large economy, has acted as a price setter, while European economies, taken separately, have, basically, acted as price takers. Hence, when real interest rates rise in the US, the respective rates in the rest of the world, and in Europe, also rise. However, when they rise in Europe the corresponding rates in the world are not necessarily affected. Next, the US being, for all intents and purposes, a closed economy is not likely to be affected by inflation expectations in Europe or elsewhere. Hence, it is rather unlikely for high inflation expectations in Europe to be transmitted into the US. Finally, the existing evidence, Chamie et al. (1994), suggests that the US cycle has led the European cycle(s) in the pre-German reunification era, a period essentially coinciding with our sample period. That correlation seems to have weakened in the nineties. In short, the US variables under consideration may be thought of as signaling variables of what is about to happen in Europe in the near future and determine inflation expectations there. From another viewpoint, Dominguez (1997) argues that monetary and exchange rate coordination agreements reached at Economic Summits, G-7, G-5, and G-3 meetings, over the period 1975 through 1993, have resulted in the US monetary policy as being the most influential among the G-3. In this regard, she identifies the US monetary dominance as arising from monetary arrangements rather than from correlated economic shocks. The preceding analysis is in disagreement with results provided by Katsimbris and Miller (1993), according to which four EMS interest rates, namely the German, Irish, Italian and the Dutch ones, Granger cause the US rate.

Our results support the hypothesis that the EMS has been functioning as a rather symmetric system, since causality between Germany and the rest of the EMS countries has been found to run in both directions (see also DeGrauwe, 1989; Katsimbris, 1993). This result, is also, in disagreement with the estimates given by authors arguing in favor of the strong version of the GDH (e.g. Karfakis and Moschos, 1990; Katsimbris and Miller, 1993). These studies suggest that the German rate causes the EMS rates, but is not caused by them. What are the reasons for such a disagreement? One might be tempted to relate them to differences in the sample sizes used. The present study uses quarterly data that extend up to 1994, whereas the previous studies use monthly data that do not extend beyond 1988.11. However, the difference between the evidence provided here and the previous results is too large to be attributed solely to sample size considerations. To that end, we have reran our systems for the period covered by the above mentioned studies and got results, tabulated in Table 6c, similar to those provided for the period 1979.2–1991.1 (Table 6b). Consequently, the evidence points again toward a dominant role for the US and a symmetric functioning in the EMS.

The most important aspect in which this study differs from the previous ones, concerns the econometric methodology adopted. In particular, most studies do not examine in a trivariate context the interest rate linkages among EMS countries, Germany and the US, but they carry out their analysis in separate bivariate frameworks. For an exception see Katsimbris and Miller (1993). As shown in Section 2, however, causality and cointegration analysis in incomplete systems may lead to totally erroneous conclusions. Moreover, when they estimate their bivariate systems in error correction form they impose specific restrictions on the error correction terms, thus implicitly imposing untested restrictions on the cointegrating vectors. Last but not least, testing causality hypotheses in cointegrated systems by means of F-tests may also lead to erroneous inference, since the F-statistics do not follow standard asymptotics, unless there is a sufficient degree of cointegration arising from the variables whose causal effects are being tested (see Section 2.1). In the present case, the evidence points towards only one cointegrating vector, thus violating the necessary condition for standard asymptotics.

4. Summary and conclusions

In this paper, we have shown the following: Firstly, necessary and sufficient conditions for the unit coefficients in individual autoregressive representations of two (or three) series to be the same with the corresponding autoregressive coefficients (α_{11}, α_{22}, α_{33}) in a VAR(1) system amount to Granger non-causality restrictions among all variables involved. Secondly, these restrictions are also sufficient for the number of unit roots in the VAR to be the same with the number of unit roots in the individual autoregressive representations of the series, i.e. for no cointegration. Thirdly, in the context of first order models with non-explosive variables, causality is also sufficient for cointegration.

Following Caporale and Pittis (1997), we have also discussed how causality and cointegration inference is affected by the omission of an important causing variable. The implications of testing for causality in incomplete systems proved to be of paramount importance in the present study.

We then reformulated the German Dominance Hypothesis within the new theoretical framework. In the bivariate VAR system three cases have been investigated: (a) the strong version of the GDH ($\alpha_{12} \neq 0$, $\alpha_{21} = 0$); (b) the symmetry hypothesis ($\alpha_{12} \neq 0$, $\alpha_{21} \neq 0$); and (c) the monetary autonomy hypothesis ($\alpha_{12} = 0$, $\alpha_{21} = 0$). Next we have allowed for the possibility that the US monetary policy could affect the EMS countries. This was done via trivariate VAR estimates in which the US interest rate was added. This has enabled us to introduce additional variations of the GDH, namely the strong, semi-strong, and weak (types 1 and 2) GDH as well as the US dominance hypothesis. These definitions take into account whether the US rate affects directly each of the individual EMS rates or indirectly via its effect on the German rate. The empirical evidence stemming from the short-term interest rate systems seems to support the symmetry hypothesis within the EMS and the US dominance hypothesis, according to which: (a) the German rate affects each of the EMS rates and is affected by them; and (b) the US rate affects the EMS rates, both directly and indirectly through its effect on the German rate.

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Appendix A: Sequential tests for parameter time invariance

To save space, the results from the tests, described below, are not reported here. They are available upon request to the authors.

(a) One-step Chow test: This is a sequence of one-step ahead Chow-type tests: We start with a minimum sample of size, say T_0, and then we augment it by one recursively, thus resulting in a sequence of F-statistics $F(1, t-k-1)$, $t = T_0 + 1....T$, given by:

$$\frac{\mathrm{RSS}_t - \mathrm{RSS}_{t-1}}{\mathrm{RSS}_{t-1}}(t-k-1)$$

(b) Break point F-tests: This is a sequence of Chow-type tests which test the model over the period 1 to $T_0 - 1$ against the whole period. A typical statistic is calculated as:

$$\frac{(\mathrm{RSS}_T - \mathrm{RSS}_{t-1})(t-k-1)}{\mathrm{RSS}_{t-1}(T-t+1)}$$

The critical values appear in the plots to be a straight line at unity. This is because the statistics reported above are scaled by one-off critical values from the F-distribution at the selected probability level. as an adjustment for changing degrees of freedom.

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