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INTEREST RATE LINKAGES WITHIN THE EUROPEAN MONETARY SYSTEM: A TIME SERIES ANALYSIS

by

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## 1. INTRODUCTION

The objective behind the foundation of the European Monetary System (EMS) in March 1979 was the creation of a system of monetary stability in Europe. It is generally accepted that the adoption of the EMS has contributed to the reduction of the intra-EMS nominal and real (bilateral and effective) exchange rate volatility relative to the pre-EMS period (Rogoff 1985, Artis 1987, and Taylor and Artis 1988). However it is not clear how much of this reduction is the result of an increased policy coordination. It has been argued that the maintainance of capital controls by the majority of the countries participating into the EMS has been a crucial element of its functioning.[1] Rogoff (1985) established that the reduction in the variability of real exchange rates in France and Italy over the EMS period was not accompanied by reduced fluctuations in real interest rate differentials and attributed this result to the presence of capital controls. The importance of capital controls is also reflected in the increased volatility of the onshore/offshore interest rate differentials (Giavazzi and Giovannini 1987, Taylor and Artis 1988). Nevertheless, it has been argued that due to "leakages", the effectiveness of controls is only temporary (Wyplosz 1988, Steinherr and De Schrevel 1988).

The higher exchange rate stability observed during the EMS period was combined with a reduction, but not the elimination, in the response of the DM cross rates to changes in the effective dollar rate (Giavazzi and Giovannini 1986, Artis 1987). Taylor and Artis (1988), on the other hand, found evidence of non-zero risk premia between DM and the other EMS currencies, but could not reject the zero risk premium hypothesis between dollar and DM. These results which highlight the asymmetric nature of the EMS are consistent with views about the dominant role of Germany in the system ("German leadership" hypothesis).

This paper addresses the issue of interest rate linkages between Germany and the other EMS countries which participate in the system's exchange rate mechanism. Firstly, it investigates whether there exist long run comovements between German and other EMS member rates by employing integration and cointegration techniques. Then, it examines whether German interest rate changes convey information about future movements of other EMS interest rates.

## 2. METHODOLOGICAL ISSUES

In a two country context with highly substitutable assets international portfolio equilibrium is specified as

$$i(t)-i(t)^{*}=E(t)s(t+1)-s(t)+v(t) \tag{1}$$

where $i(t)$ and $i(t)^{*}$ denote domestic and foreign interest rates respectively, $E(t)$ refers to expectations at time $t$, $s(t)$ is the logarithm of the exchange rate (domestic currency per foreign currency unit) and $v(t)$ represents the risk premium. The existence of risk premia implies a departure from uncovered interest parity. If the expected exchange rate change and the risk premium are both stationary, then the interest rate differential is also stationary, in which case there exists a long run equilibrium relationship between $i(t)$ and $i(t)^{*}$.

In a more general context, the existence of a long run interest rate relationship can be looked at from the viewpoint of cointegration (Engle and Granger 1987) by investigating for stationarity the process

$$i(t)-bi(t)^{*}=\omega(t) \tag{2}$$

where, in contrast to (1), b is allowed to deviate from unity and the right hand side of (1) is supressed into $\omega(t)$. The reason for allowing $b\neq1$ can be clarified by considering interest income taxation and/or the possibility of measurement errors.

The lack of a harmonized tax system within the EMS and the existence of interest income taxation modifies the interest parity condition to:

$$i(t)[1-\tau]-i(t)^{*}[1-\tau^{*}]=z(t) \tag{3}$$

where $\tau$ and $\tau^{*}$ are the rates at which interest income is taxed. Equation (3) is equivalent to (2) with $b = [(1-\tau^{*})/(1-\tau)]$ and $\omega(t) = [z(t)/(1-\tau)]$. Clearly if $\tau\neq\tau^{*}$, then $b\neq1$. Moreover, in the presence of inter-country differences in the maturity and definition of assets, deviations of b from unity can be attributed to measurement errors in published series. Assume that the interest parity condition (1) holds for some theoretical variables $r(t)$ and $r(t)^*$ which are measured as

$$i(t)=\beta+\gamma r(t)+u(t) \tag{4}$$

$$i(t)^*=\beta^*+\gamma^* r(t)^*+u(t)^* \tag{5}$$

where $\beta$, $\gamma$, $\beta^*$ and $\gamma^*$ capture systematic measurement errors and $u(t)$ and $u(t)^*$ are stationary random errors. Substituting for $r(t)$ and $r(t)^*$ into (1), we end up with (2), where $b = \gamma\gamma^{*-1}$, and $\omega(t) = \gamma[E(t)s(t+1)-s(t)+v(t)]+u(t)-\gamma\gamma^{*-1}u(t)^*$.

In applying the cointegration testing procedure, firstly the hypothesis that the individual series contain a unit root, i.e. they are integrated of order one (I(1)), needs to be tested, since in this case standard statistical inference is invalidated (Phillips 1986). For this end the Dickey-Fuller (DF) and augmented DF (ADF) tests (Fuller 1976, Dickey and Fuller 1981) can be used.[2] Recently Perron (1986), Phillips (1987), Phillips and Perron (1988) proposed non-parametric tests which allow for weakly dependent and heterogenous disturbances. If $i(t)$ and $i(t)^*$ both contain one unit root, then a test of cointegration can be obtained as the DF (ADF) statistic of the residuals of regression (2) or alternatively as the DW statistic (CIRDW) of (2).

Engle and Granger (1987) have shown that if a set of variables are cointegrated then there exists an error



correction (EC) representation and vice versa, in which case there must be Granger-causality (Granger 1969) in at least one direction. It follows that vector autoregressive (VAR) models in differences lead to misspecification if the variables are cointegrated.

## 3. EMPIRICAL RESULTS

The empirical analysis is carried out using monthly data on short term domestic nominal interest rates for Belgium (B), France (F), Germany (G), Ireland (IRL), Italy (I) and the Netherlands (NL) over the period April 1979 to November 1988.[3] The DF and Phillips and Perron (1988) test statistics of the null hypothesis of one unit root as well as the joint hypothesis of a unit root and a zero mean for both the levels and the first differences are reported in Table 1. The hypothesis of a unit root in the levels of interest rates can not be rejected for any of the series concerned, regardless of the statistic and the value of the truncation lag (l=1,5) used. By contrast, the hypothesis of a unit root in the first differences is rejected in all cases with the exception of Belgium, for which the joint hypothesis is not rejected. The non-rejection of the joint hypothesis is attributed to the presence of a non-zero mean.[4] These results suggest that all interest rate series are I(1).

The tests for cointegration between German and each of the other EMS interest rates, which are reported in Table 2, fail to reject the null hypothesis of noncointegration at the 0.05 significance level in all cases. In view of these results, which indicate that the EMS has not been dominated by stationarities in exchange rate changes and/or in risk premia, the investigation of the hypothesis that German interest rates are indicators of future movements of other EMS member rates and not vice-versa is carried out by means of VAR models in first differences.

The selection of the lag order of the VAR systems is based on two alternative and well established criteria: first, the likelihood ratio test corrected for the degrees of freedom (Sims 1980) and second, Akaike's (1969) information criterion based on the final prediction error (FPE) as implemented by Hsiao (1979). In applying these criteria, the monthly frequency of the data favored the selection of 12 as the maximum lag length.

The lags selected on the basis of each criterion are reported in Table 3. The differences between these lag structures stem from the feature that, contrary to Sims'(1980)-criterion which results in unrestricted VAR systems, FPE does not impose a uniform lag length on all the variables. Efficient estimates of these restricted VAR systems can be obtained by GLS.

The results of Granger-causality F-tests of the unrestricted VAR systems which are given in Table 3 suggest that the past history of changes in German rates has a predictive power for interest rate changes in other



EMS members with the exception of Ireland. On the other hand, German interest rate movements seem to be Granger-exogenous with respect to other EMS member rates.

In order to check the stability of the equations in which the German interest rate had a significant information content, predictive failure tests over the last six months of the sample were carried out. In all cases the tests did not reject the hypothesis of no predictive failure. Furthermore, to investigate the sensitivity of these equations with respect to exchange rate realignments, predictive failure tests were conducted over the period following the major realignment of 21st March, 1983. Again no evidence of predictive failure was found.[5]

The GLS results of the restricted VAR systems in Table 3, replicate the conclusions obtained from the unrestricted OLS estimates. This similarity in the findings indicates that the causality results are insensitive to the lag specification.

On the whole, the results suggest that the Belgian, French, Italian and Dutch interest rate changes can be predicted using information on the past evolution of German interest rate changes, but not conversely. This evidence seems to be consistent with arguments about the pivotal role of Germany in the EMS. It is worth noting that the results are similar across countries with substantial differences in the degree of capital mobility.

To test for instantaneous causality between German and other EMS member rates, the contemporaneous innovation correlations for each bivariate system has been calculated. The results of Table 4 yield significant correlations with respect to Belgium, Italy and the Netherlands. In the light of the EMS experience, it seems plausible to interpret the instantaneous causality results on the basis of the arguments pertaining to the German leadership in the EMS. In the case of France the lack of contemporaneous causality might be possibly attributed to the nature of capital controls.

**4. CONCLUSIONS**

This paper has concentrated on the bivariate analysis of interest rate linkages within the EMS. Cointegration tests have not revealed the existence of systematic interest rate relationships in the long run between Germany and any of the other EMS countries. This finding may be attributed to the non-stationarity of either the expected exchange rate movements or the risk premia. Bivariate vector autoregression analysis has been used to investigate the information content of German interest rates about the future course of other member rates. An interesting aspect which highlights the dominant role of Germany in the EMS is the evidence of unidirectional interest rate linkages from Germany to other EMS countries, regardless of the nature and degree of restrictions on capital movements.



NOTES

1. Denmark, France, Ireland and Italy have been using capital controls of various degrees of severety, whereas Belgium has adopted a dual exchange rate as a substitute for controls (Steinherr and De Schrevel 1988).

2. The distribution of the DF or ADF test is not invariant with respect to the presence of a constant or a deterministic trend. Dickey, Bell and Miller (1986) recommend against the inclusion of a trend since this would make a random walk model look stationary. However, they suggest the inclusion of a constant.

3. The data have been obtained from OECD, Main Economic Indicators. These are yields on three-month (B,IRL) or six-month (I) Treasury bills and on three-month loans (F,G,NL). Due to data limitations, the French series cover the period 1981:5-1988:11, and Denmark has not been included into the analysis.

4. The statistic $\tau_{\alpha\mu}$ (Dickey and Fuller, 1981) is 0.16 and therefore isignificant, suggesting the presence of a zero mean.

5. The test statistic has an approximate F distribution (Chow 1960). Over the period May-November 1988 the F-values for B, F, I and NL were $F(6,73)=0.30$, $F(6,81)=0.49$, $F(6,100)=0.28$ and $F(6,73)=0.73$ respectively. Over the post-March 1983 period the corresponding values were $F(68,11)=1.68$, $F(68,19)=0.18$, $F(68,38)=0.37$ and $F(68,11)=1.11$.

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TABLE 1
Root Tests for Short Term Domestic Interest Rates

| | Country | | | | | |
|---|---|---|---|---|---|---|
| Statistic | B | F | G | IRL | I | NL |
| | | | LEVELS | | | |
| | | | Dickey - Fuller | | | |
| $\tau_\mu$ | -1.15 | -1.65 | -0.87 | -1.18 | -0.92 | -1.14 |
| $\Phi_1$ | 0.68 | 0.93 | 0.39 | 0.49 | 0.43 | 0.61 |
| | | | Phillips - Perron | | | |
| l=1 | | | | | | |
| $Z(t_{\alpha}^{*})$ | -1.46 | -1.58 | -1.16 | -1.71 | -0.77 | -1.26 |
| $Z(\Phi_1)$ | 1.01 | 0.07 | 0.69 | 1.33 | 0.30 | 0.81 |
| l=5 | | | | | | |
| $Z(t_{\alpha}^{*})$ | -1.62 | -1.91 | -1.44 | -1.63 | -1.16 | -1.55 |
| $Z(\Phi_1)$ | 1.36 | 1.01 | 1.04 | 1.19 | 0.69 | 1.22 |
| | | | FIRST DIFFERENCES | | | |
| | | | Dickey - Fuller | | | |
| $\tau_\mu$ | -8.21 | -12.87 | -7.50 | -7.84 | -9.15 | -9.77 |
| $\Phi_1$ | 3.67 | 82.72 | 28.11 | 30.80 | 75.99 | 48.01 |
| | | | Phillips - Perron | | | |
| l=1 | | | | | | |
| $Z(t_{\alpha}^{*})$ | -8.18 | -11.89 | -7.56 | -7.91 | -8.81 | -9.76 |
| $Z(\Phi_1)$ | 2.70 | 79.28 | 28.60 | 31.39 | 73.54 | 47.88 |
| l=5 | | | | | | |
| $Z(t_{\alpha}^{*})$ | -8.22 | -12.04 | -8.29 | -7.66 | -10.18 | -9.96 |
| $Z(\Phi_1)$ | 4.12 | 79.58 | 30.02 | 29.32 | 69.18 | 49.94 |

NOTES: $\tau_\mu$ and $Z(t_{\alpha}^{*})$ have a critical value of -2.89 (Fuller 1976) and $\Phi_1$ and $Z(\Phi_1)$ a corresponding value of 4.71 (Dickey and Fuller 1981) at the 0.05 level for T=100.

TABLE 2
Tests of Cointegration Between German and Other EMS Interest Rates

| EMS Country | DF | ADF(4) | CIRDW |
|---|---|---|---|
| Belgium | -2.98 | -2.26 | 0.29 |
| France | -2.72 | -1.64 | 0.21 |
| Ireland | -2.28 | -2.86 | 0.22 |
| Italy | -1.61 | -1.93 | 0.08 |
| Netherlands | -3.03 | -2.87 | 0.28 |

NOTE: ADF(4) denotes that four lags of the dependent variable are included in the ADF test. The critical values of the DF, ADF(4) and CIRDW tests of cointegration are -3.37, -3.17 and 0.39 respectively at the 0.05 level for T=100 (Engle and Yoo 1987).

TABLE 3
Granger-Causality Tests of Unrestricted and Restricted VAR Systems

| System | Dep. V/ble | Unrestricted System: Lags | Unrestricted System: MSL of F-test | Restricted System: (m,n) | Restricted System: MSL of $X^2$-test |
|---|---|---|---|---|---|
| B,G | B | 12 | 0.000* | (11,12) | 0.000* |
| | G | 12 | 0.478 | (3,1) | 0.906 |
| F,G | F | 1 | 0.014* | (12,5) | 0.046* |
| | G | 1 | 0.704 | (1,5) | 0.087 |
| IRL,G | IRL | 6 | 0.294 | (8,1) | 0.220 |
| | G | 6 | 0.683 | (3,1) | 0.924 |
| I,G | I | 3 | 0.002* | (11,3) | 0.007* |
| | G | 3 | 0.752 | (3,1) | 0.810 |
| NL,G | NL | 12 | 0.012* | (9,5) | 0.044* |
| | G | 12 | 0.070 | (3,1) | 0.177 |

NOTE: (m,n) denotes the number of lagged dependent and independent variables respectively. The F and $X^2$ tests are calculated for the joint hypothesis that the coefficients of the lags of the independent variable are zero. MSL=marginal significance levels. An asterisk (*) indicates significance at 5%.

TABLE 4
Correlation Coefficients of Innovations in VAR Systems

| Estimation Method | System: (G,B) | (G,F) | (G,IRL) | (G,I) | (G,NL) |
|---|---|---|---|---|---|
| OLS | 0.36* | 0.06 | -0.09 | 0.25* | 0.67* |
| GLS | 0.40* | 0.03 | -0.03 | 0.36* | 0.63* |

NOTE: * indicates significance at 5%.

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