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Cointegration, Uncovered Interest Parity and the Term Structure of Interest Rates: Some International Evidence

Dimitris A. Georgoutsos[1]
Athens University of Economics and Business
and
Georgios P. Kouretas[*]
University of Crete

Abstract

This paper addresses the issue of the empirical investigation of monetary policy independence as this is manifested in the inter–relationships between domestic and foreign money market interest rates. Instead of following an ad–hoc econometric approach, we have imposed a specific economic structure on the proposed model by establishing a link of the yield curves of two different countries through the Uncovered Interest Rate Parity, UIP. The expectations hypothesis of the term structure and the UIP imply certain overidentifying restrictions on the cointegrating space of a vector autoregressive process consisting of the interest rates of the two markets. The model has been tested on data from the domestic US money market and the euromark and euroyen markets. The main finding of our analysis is that we reject the overidentifying restrictions of the models for the USD/DEM case but we are unable to reject them for the USD/JPY case, at the one percent significance level and this im-

[1] Dimitris A. Georgoutsos, Department of International and European Economics, Athens University of Economics and Business, 76 Patission Street, GR-10434, Athens, Greece, tel. 003018203441, fax. 003018214122 email: D.Georgoutsos@aueb.gr; Georgios P. Kouretas, Department of Economics, University of Crete, University Campus, GR-74100, Rethymno, Greece, tel. 003083177412, fax. 003083177406, email: kouretas@econ.soc.uoc.gr.

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plies that the term spreads of the euroyen market are being affected, in the long run, by changes of the US short Fed funds rate.

Keywords: cointegration, expectations theory, uncovered interest parity, eurocurrency markets, Granger causality
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1. Introduction

The purpose of this paper is to bring together two of the most important branches of modern finance literature, the expectations hypothesis (EH) of the term structure of interest rates and the Uncovered Interest Rate Parity (UIP). By employing the framework provided by cointegration theory it is shown that the satisfaction of the necessary conditions for the EH in one country when combined with the UIP hypothesis imply that the term spread in the second country must be stationary.

The implications of this result are evident. When we focus on the money market term structure, the common trend under the EH is usually identified with the interest rate under the control of monetary authorities. This common trend can be the funds rate in the USA, the Lombard rate in Germany, the very short-term rate on repurchase agreements in France and Italy or the overnight rate in other countries (Estrella and Mishkin, 1997). Furthermore, it is shown that the rates of corresponding maturities between two countries are cointegrated then the term structure in one country must be directly dependent on the other. Dependence here is used in the sense that the deviation of the rates can not be "permanent" and that a mean reversion must take place in the longer term. This approach offers an initial test to problems studying the independence of monetary policies in different countries. A side effect of these tests also shows that past values of the term premiums in one market and of the spreads of rates denominated in different currencies are optimal predictors of the term premium changes in the other country. Moreover, there must exist a Granger-type causal relationship from the term premium of one country to the short rate changes of the other and not the other way around.

The exposition of some interesting cases can make the above testing procedure clearer. In the case that the EH can not be rejected in the "major" economy but the UIP condition is not validated by the data then changes in the monetary policy can not be transmitted, in a permanent way, to the other economy. An even more interesting case is when EH is satisfied in one country but the UIP holds only for short-term maturities. In this situation changes in monetary policy are transmitted to longer-term maturities only in the first country while the longer-term rates of the second appear to be immunized. To the extent that the real economy is affected by changes of the longer-term rates, and vise-versa, this immunization signifies that other factors like the inflation rate or the growth rate of the economy might be important determinants of those rates.

If a country was isolated then its term structure spread would be strictly dependent on underlying factors such as its central bank policies and expectations by local investors on future real activity and inflation. However, since capital flows are transmitted across countries the same factors that influence a country's own term premium may influence term premiums of other countries, e.g. inflationary expectations (since economic and inflationary conditions are interrelated across countries). Also, if monetary policy is used to influence the term structure it must be taken into account for the influence of foreign term structures. A simple test would imply the existence of a causal relationship between the two term premiums or if they are non-stationary variables the presence of a common trend. This is the approach taken, for example, by Madura *et al*. (1998) in which they search for cointegrating relationships among the term premiums of different countries. We depart from this approach in the present paper by establishing the link for the above-mentioned relationship to exist. This link is provided by the

UIP hypothesis that is a necessary condition to hold for the term premiums to be interrelated in the long run as this is defined in cointegration theory.

The question of whether various money markets are segmented has been addressed in many previous studies. However, the major part of this literature is concerned with the interest rate linkages between assets traded in different markets but are denominated in the same currency. Thus, Swanson (1988) had shown that eurodollar rates adjust faster to the domestic market changes than the other way around while contemporaneous causality prevails when weekly data were used. Fung and Isberg (1992) concluded that there existed causalities between the US and the Eurodollar markets from both directions depending on the period examined. Bloocha-oom and Stansell (1990) established that dollar denominated interest rates in Hong-Kong and Singapore markets reacted instantaneously to changes in the US markets. On the other hand previous studies, e.g. Levin (1974), had provided support for the existence of segmentation between the US and the Eurodollar market. An exception to the previous literature is the study by Kasman and Pigott (1988) who concluded that the divergences between interest rates across countries have been accentuated since the inception of floating exchange rates in 1973. Finally, Madura *et al*. (1998) have established a long-run equilibrium relationship among the term premiums of various countries and that the forecasting ability is enhanced by the incorporation of error correction terms.

The model we developed was tested using monthly data from the domestic USD and the euromark and euroyen money markets. The main finding is that we failed to reject the overidentifying restrictions of the model, at the 1% significance level, for the USD/JPY case while we rejected them for the USD/DEM case. The failure in this last case was attributed to the lack of empirical support for the Uncovered Interest Parity condition. The weak support to the model was also reconfirmed when causality tests were conducted between the interest spreads and short rate changes.

The rest of the paper is organized as follows. In section 2 the expectations hypothesis of the term structure and the uncovered interest rate parity condition are presented and then the long run exclusion restrictions in a cointegrated Vector Autoregressive (VAR) are derived. In section 3, the econometric methodology is presented. Section 4 presents and discusses the empirical results with section 5 providing our concluding remarks.

2. Term structure, uncovered interest parity and cointegration analysis

Let R(t, n) and R(t,1) denote the n and 1-period rates of interest respectively at time t. The risk adjusted expectations theory then states that the long interest rate is the average of the current one period rate and the expected, as at time t, future one period rates plus a premium reflecting risk and / or liquidity considerations; that is,

$$R(t,n) = (1/n)\sum_{j=1}^{n} E_t R(t+j-1,1) + P(t,n) \tag{1}$$

where $P(t,n) = (1/n)\sum_{j=1}^{n} p(j,t)$ represents the premium component.

A more enlightening version of eq. (1) can be obtained if we subtract *R(t,1)* from both sides of the equation and rearrange it to give:

$$S(t,n) = \sum_{j=1}^{n-1} (1 - j/n) E_t \, \Delta R(t+j,1) + P(t,n) \,, \tag{2}$$

where

$$S(t,n) = R(t,n) - R(t,1) \text{ and } \Delta R(t+j,1) = R(t+j,1) - R(t+j-1,1).$$

According to eq. (2) the spread between long and short interest rates will be equal to a weighted sum of the expected, at time t, changes of the short interest rate plus the risk premium part. The weighting scheme implies that expected short term interest rate changes in the near future carry more weight in determining the spread than do expected short rate changes in the more distant future.

If the nominal interest rates are integrated I(1) processes then an interesting testing implication for the expectations theory can be derived. Assuming that the first differences of nominal interest rates and the premia are stationary variables then $R(t, n)$ and $R(t,1)$ must be cointegrated with a cointegrating vector (1, -1). Since this implication of the model applies to any "long" interest rate $R(m, t)$, $m \neq n$, then if we consider a set of n interest rates we expect not to be able to reject the hypothesis that the $(n - 1)$, n dimensional linearly independent cointegrating vectors, (1, -1, 0,.,....,0), (1, 0, -1, 0,..0) .... ,(1, 0, 0,.....-1), form the basis of the cointegration space.

A second implication of eq. (2) is that the spread between long and short rates is an optimum forecast of future short interest rates changes. This suggests that the spread must Granger - cause the short rate changes while causality of the opposite direction should not exist.

On the other hand, the uncovered interest rate parity implies that the yields of domestic and foreign assets can differ only by the expected change in the price of foreign exchange, which in a formal representation can be written as:

$$R(t,n) = R_f(t,n) + (1/n)(E_t s(t+n)/s(t) - 1) + D(t) , \qquad (3)$$

where $R_f(t,n)$ refers to the return on the foreign asset, $s(t)$ is the exchange rate and $D(t)$ is a country specific risk premium.

If the implications of the afore-mentioned theories cannot be rejected then an interesting result is derived for the term structure of the foreign interest rates. If one substitutes eq. (3) in (2) then it can be easily shown that the spread on the foreign interest rates is given by:

$$S_f(t,n) = R(t,n) - R(t,1) - (1/n)(E_t s(t+n)/s(t) - 1) + E_t(s(t+1)/s(t) - 1) \qquad (4)$$

where $S_f(t,n) = R_f(t, n) - R_f(t, 1)$. According to equation (4) if the term structure applies to the domestic country and the exchange rate percentage changes are *I(0)* variables, then the spread on foreign interest rates is an *I(0)* variable itself i.e. the term structure applies to foreign interest rates as well.

Cointegration implied by the above considerations is of a very special type. Specifically, if both theories hold then if we consider a set of n domestic and n foreign interest rates then the cointegration space must have a rank of (2n-1) and the following set of identifying restrictions:

$$\begin{array}{cccccccccc} R(t,1) & R(t,2) & \ldots & \ldots & R(t,n) & R_f(t,1) & R_f(t,2) & \ldots & \ldots & R_f(t,n) \end{array}$$

$$\begin{bmatrix} 1 & -1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 1 & 0 & -1 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ .. & .. & .. & .. & .. & .. & .. & .. & .. & .. \\ 1 & 0 & 0 & 0 & -1 & 0 & 0 & 0 & 0 & 0 \\ 1 & 0 & 0 & 0 & 0 & -1 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 & -1 & 0 & 0 & 0 \\ .. & .. & .. & .. & .. & .. & .. & .. & .. & .. \\ 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 & -1 \end{bmatrix} \qquad (5)$$

should not be rejected.

Another, interesting testing implication of equation (4) is that the difference of the domestic spread between long and short rates from the foreign one, is an optimum forecast of future changes of the exchange rate. Furthermore, one should be able to reject that any other variable, e.g. changes in short interest rates, has an explanatory power over the difference of the two spreads.

Finally, forecasts of interest rate changes can be enhanced from the inclusion in the auto-regressive process of an error correction term relating to the cointegrating vectors. Variables that would be found weakly exogenous could be candidates for the common stochastic trends that drive the system.

Under the testing framework given above we can discern some interesting alternative scenarios:

CASE 1: (presented above) one common stochastic trend possibly related to the short term USA interest rates (policy instrument). This would be evidence of total lack of an independent monetary policy.

CASE 2: the expectations hypothesis applies only to one country (e.g. US) and the UIP does not hold (the identifying restrictions refer to the first (n-1) lines of the matrix above). This is evidence of an independent monetary policy where in only one of the two countries the monetary authorities have control over the shape of the yield curve (in the sense that one common trend is responsible for the determination of the yields).

CASE 3: the expectations hypothesis holds for both countries but UIP does not hold. In this case 2x(n-1) restrictions should be satisfied (the restrictions on R above apply also for the Rf rates). This is also evidence of an independent monetary policy where two common trends, one for each country, drive the system.

3. Econometric methodology

Our cointegration analysis is based on the multivariate cointegration technique developed by Johansen (1988, 1991) and extended by Johansen and Juselius (1990, 1992) which is a Full Information Maximum Likelihood estimation method.

Consider a $\boldsymbol{p}$-dimensional vector time series z_t with an autoregressive representation which in its error correction form is given by

$$\Delta z_t = \Gamma_1 \Delta z_{t-1} + ... + \Gamma_{k-1} z_{t-k+1} + \Pi z_{t-k} + \gamma D_t + \mu + \varepsilon_t \qquad (6)$$

where z_t is vector of stochastic variables, $\varepsilon_t \sim Niid_p(0,\Sigma)$. The parameters $(\Gamma_1,.........,\Gamma_{k-1},\gamma)$ define the short-run adjustment to the changes of the process, whereas $\Pi = \alpha\beta'$ defines the short-run adjustment, α, to the cointegrating relationships, β. D_t is a vector of nonstochastic variables, such as centered seasonal dummies which sum to zero over a full year by construction and are necessary to account for short-run effects which could otherwise violate the Gaussian assumption, and/or intervention dummies; μ is a drift and T is the sample size.

Model (6) will be treated as a benchmark model within which all the subsequent hypotheses are tested. Since the parameter set $\theta = (\Gamma_1, \ldots\ldots, \Gamma_{k-1,} \Pi, \gamma, \mu, \Sigma)$ varies unrestrictedly, it follows that the I(1) model is a submodel of (6). In the unrestricted form, therefore, model (6) corresponds to the I(0) model. In the statistical sense the I(0) model is the most general, since the higher-order models are nested in the model.

Johansen (1991) shows that if $Z_t \sim I(1)$, the following restrictions on model (6) have to be satisfied:

$$\Pi = \alpha\beta' \tag{7}$$

where Π has reduced rank, r, α and β are (pxr) matrices, and

$$\Psi = \alpha_{\perp}(-I + \Gamma_1)\beta_{\perp} = \varphi\eta' \tag{8}$$

where Ψ is a (p-r)x(p-r) matrix of full rank, φ and η are (p-r)x(p-r) matrices, and $\alpha_{\perp}$ and $\beta_{\perp}$ are px(p-r) matrices orthogonal to α and β, respectively. The parameterization in (7) and (8) facilitates the investigation of, on the one hand, the r linearly-independent stationary relations between the levels of the variables and, on the other hand, the p-r linearly-independent non-stationary relations. This duality between the stationary relations and the non-stationary common trends is very useful for a full understanding of the generating mechanisms behind the chosen data.

4. Empirical results

The vector autoregressive representation model, as defined in (6) above, has been estimated on one, three and twelve months interest rates. The calculations of all tests have been performed using the program CATS 1.1. in RATS 4.30, Estima Inc. The data is monthly and cover the period May 1991 to May 2001 and they refer to the Fed Funds Rate, and the Libor offer rates for the Deutschemark and JPY respectively. The data has been obtained from the Bloomberg Financial Services databank (figures 1 and 2). Two bilateral models have been estimated, US/Germany, and US/ Japan.

Before proceeding with the presentation of our results and for comparison purposes it would be interesting to discuss recent evidence from empirical tests on the EH and the UIP condition. The international evidence on the EH is rather puzzling. When we look at the returns of securities with maturity of one year or less the EH is more often accepted when European data are studied. Gerlach and Smets (1997) have shown that euro-rates term spreads contain information on future short-term interest rates for 17 countries with the weakest support coming from eurodollar rates. Dahlquist and Jonsson (1995) fail to reject the EH for Swedish Treasury Bill rates while Hall, Anderson and Cranger, (1992) find supportive evidence to the *EH* coming from data on Treasury Bill rates. Concerning the returns on securities with maturity of more than one year the evidence is again most favorable to non-US data e.g. Jorion and Mishkin (1991), Hardouvelis (1994). The evidence on the UIP is even more perplexing. In an early study Cumby and Obstfeld (1984), have shown that under rational expectations the errors are non-stationary. Other authors, Fama (1984), invoke the assumption that the forward exchange rate is an unbiased predictor of the future exchange rate and have tested for the Covered Interest Rate Parity. McCallum (1994) showed that testing for those two conditions is not equivalent while Johansen and Juselius (1992) find supportive evidence for the UIP when it is tested jointly with the Purchasing Power Parity condition.

The statistical tests rely upon the Gaussian assumption of the error terms in (6). Therefore, in Table 1 we present residual misspecification tests for the two cases of the VAR model with 4 lags. We note that our conditional models are well specified with respect to the conditional heteroscedasticity of the residuals since the ARCH statistic was never found to be statistically significant. A problem was detected however, in the USD/DEM case, concerning the normality hypothesis of the error terms. In three out of the six estimated equations the NORM statistic has been found to be significant. This is attributed to the kurtotic behavior of the error terms, as the η_4 statistic indicates. This evidence is not weakening our cointegration analysis since it is well documented in the literature that the test statistics are inaccurate only when the *fat–tails* behavior of the residuals is attributed to skewness. Furthemore, we present evidence on multivariate autocorrelation and all our tests have rejected the null hypothesis of serial correlation. Finally, we reject the null hypothesis of multivariate normality but this is not worrying since it is due to the presence of kurtosis rather than skewness in the residuals, (Gonzalo, 1994).

In Table 2 we provide statistical justification for the presence of the selected variables in the VAR model. Thus, we reject the null hypothesis that any of the variables should be excluded from the long run equilibrium relationship. Moreover, in only one case, the twelve months DEM rate, we were unable to reject the null hypothesis that the variable is weakly exogenous to the long-run parameters of the system in the sense that there is no explanatory power of a linear combination of the term premiums on the 12-months rate changes. Finally, we are well justified in employing the framework of cointegration theory since all our variables exhibit a non-stationary behavior.

Table 3 presents the evidence for the existence of a cointegrating behavior among the variables. The trace test has been calculated for all possible values of the rank of the cointegration space, r. The 5% critical values represent the case where a constant is present but it is restricted to lie within the cointegration space

(MacKinnon *et al.*, 1999, Table III). The results offer a strong validation of our model of section 2. If both the EH and the UIP hold then we expect to find (2n–1) cointegrating vectors. In our case the presence of four against five cointegrating vectors is easily rejected for both of the models. This evidence seems to imply that a single common stochastic trend drives the entire system in the long run.

In addition to the formal test, Juselius (1995) suggests that the results from the trace test statistics should be interpreted with some caution for two reasons. First, the conditioning on intervention dummies and weakly exogenous variables is likely to change the asymptotic distributions to some unknown extent. Second, and more relevant for our case, the asymptotic critical values may not be very close approximations in small samples. Thus, Juselius (1995) suggests the use of the additional information contained in the roots of the characteristic polynomial. In Table 3 we also report the modulus of the six largest roots of the companion matrix for both the unrestricted and the restricted cointegration space cases. According to the theoretical exposition we expect to have one root almost equal to one and the others well below it. The evidence is encouraging since only one root is above 0.95 while the others are close to 0.90 and below it.

Table 4 provides test statistics based on the overidentifying restrictions implied by the theoretical model. Johansen and Juselius (1994) developed a likelihood ratio statistic that is distributed as χ^2 with $\nu=\Sigma_i(p-r+1-s_i)$ where p stands for the number of variables, r for the number of cointegrating vectors and s_i for the number of freely estimated variables in vector i. The results show that the overidentifying restrictions are rejected for the USD/DEM case while we are unable to reject them for the USD/JPY case at the 1% significance level. We have tried then to identify the reasons of failure for the USD/DEM case and therefore we imposed separately the restriction implied by the EH and the UIP conditions. In case 2 the EH for the Fed Funds market is easily rejected while we fail to reject the case for the DEM rates at the 3% significance level (case 4). Furthermore, the restrictions referring to the UIP condition are easily rejected as well. These results confirm previous ones in the literature that indicate weak support for the EH for the non-US dollar market rates. The supportive evidence in the first case for the USD/JPY rates probably indicates that the presence of the JPY rates enhances the informative power of the model.

Since weak evidence in favor of the model was found for the USD/JPY case, we decided to test an additional implication of our model, i.e. that the spread is an optimal predictor of future short rate changes while the opposite direction of the causality should not hold. An interesting implication of this model is that the above testable hypothesis holds also for changes of the short rate of the "foreign" currency. In Table 5 we present the evidence from the Granger-causality tests. In the euroyen market the null hypothesis that there is no causality, is rejected for the one to three month case under both possible directions of the causality. However, the theoretical implications are validated when the one and the twelve-month interest rates are examined. In this case we fail to reject the null that changes in short rates do not cause the spread of these two rates. We then applied the same tests when the short rate is the one-month FED rate. Contrary to the previous results, the evidence supports the theoretical results when the one and three-month JPY rates are examined in conjunction with the one month FED rate.

5. Concluding remarks

This paper develops a testing methodology to analyze the issue of independence of the monetary policy as this is reflected on the yield curves of the "domestic" and the "foreign" economy. The usual testing approach to this problem has been concerned with the existence of a relationship among the interest rates of corresponding maturities between the two countries. In this paper we take a more “structural” approach which establishes a channel between the “domestic” and “foreign” interest rates through the existence of the Uncovered Interest Rate Parity condition. Within the framework provided by cointegration theory, the restrictions in the cointegration space that must be satisfied for the expectations theory of the term structure and the Uncovered Interest Parity to hold have been derived. The major implication of this model is that changes in monetary policy, through the short interest rates, in the “foreign” country are transmitted to the domestic country’s term premium.

The model has been tested on monthly data from the domestic US dollar and the euromark and euroyen money markets. We failed to reject the overidentifying restrictions of the model, at the 1% significance level, for the USD/JPY case while we rejected them for the USD/DEM case. The failure in this last case was attributed to the lack of empirical support for the Uncovered Interest Parity condition. The weak support to the model was also reconfirmed when causality tests were conducted between the interest spreads and short rate changes.

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Table 1: *Residual misspecification tests of the model with k = 4*

USD –DEM

Eq.	σ_ε	ARCH(4)	η_3	η_4	NORM(4)	R^2
FED1	0.15	0.35	0.02	4.27	10.0*	0.68

FED3	0.16	3.05	−0.07	3.48	2.66	0.52
FED12	0.22	4.62	0.11	2.98	0.44	0.38
DEM1	0.15	3.9	−0.74	5.31	14.9*	0.57
DEM3	0.15	1.05	−1.63	8.57	33.3*	0.37
DEM12	0.18	0.98	−1.20	7.93	27.6*	0.30

USD–JPY

Eq.	σ_ε	ARCH(4)	η_3	η_4	NORM(4)	R^2
FED1	0.15	0.80	−0.29	3.78	4.95	0.69
FED3	0.16	1.54	−0.48	4.34	8.90	0.49
FED12	0.22	4.94	−0.17	2.73	0.76	0.38
JPY1	0.15	3.52	0.26	3.47	2.90	0.55
JPY3	0.13	1.29	0.24	3.57	3.41	0.41
JPY12	0.15	3.67	0.15	3.42	2.62	0.35

*Notes: σ_ε is the standard error of the residuals, η_3 and η_4 are the skewness and kurtosis statistics. ARCH is the test for heteroscedastic residuals, and NORM the Jarque–Bera test for normality. The ARCH and NORM statistics are distributed as χ^2 with 4 and 2 degrees of freedom, respectively and the LB statistic is distributed as χ^2 with 36 degrees of freedom. *(**) denotes significance at the 5% (1%) level.*

Multivariate Residual Diagnostics

Case	L–B(29)	LM(1)	LM(4)	χ^2 (12)
USD–DEM	0.56	0.10	0.67	0.00
USD–JPY	0.41	0.75	0.97	0.00

Notes: L–B is a multivariate version of the Ljung–Box test statistic for residual autocorrelation based on the estimated auto– and cross– correlations of the first [T/4=64] lags. LM(1) and LM(4) are tests for first and fourth order autocorrelation distributed as χ^2 with 49 degrees of freedom. and χ^2 with 12 degrees of freedom is a multivariate version of the Shenton–Bowman test for normality. Numbers reported are marginal significance levels.

Table 2: *Tests for long–run exclusion, stationarity and weak exogeneity*

	L-R exclusion		Stationarity		Weak exogeneity	
Variable	**USD–DEM**	**USD–JPY**	**USD–DEM**	**USD–JPY**	**USD–DEM**	**USD–JPY**
FED1	18.1*	39.9*	10.45*	41.52*	17.43*	33.06*
FED3	16.9*	40.5*	10.43*	41.38*	15.6*	22.49*
FED12	29.8*	35.7*	10.31*	40.95*	13.2*	12.00*
DEM1 (JPY1)	17.9*	29.0*	7.92*	18.44*	17.21*	28.43*
DEM3 (JPY3)	18.7*	23.9*	7.95*	18.86*	16.33*	20.15*
DEM12 (JPY12)	20.2*	10.3*	8.03*	20.49*	7.89	12.47*

Notes: The long–run exclusion restriction and the weak exogeneity tests are likelihood ratio tests distributed as χ^2 with five degrees of freedom. and the 5% critical value is 11.07. The stationarity test is also a likelihood ratio test distributed as χ^2 with two degrees of freedom and the 5% critical value is 5.99.

Table 3: *Testing the Rank of the I(1) Model*

H0:r	USD–DEM	USD–JPY	TRACE 5%
0	172.7*	181.9*	103.84
1	127.8*	131.6*	76.96
2	83.6*	85.7*	54.09
3	46.7*	46.5*	35.19
4	20.3*	21.5*	20.25
5	4.65	8.87	9.17

Notes: p is the number of variables, r is the rank of the cointegration space. The 5% critical values are taken from MacKinnon et al. (1999, Table III). For each case a structure of four lags was chosen according to a likelihood ratio test, corrected for the degrees of freedom (Sims, 1980) and the Ljung–Box Q statistic for detecting serial correlation in the residuals of the equations of the VAR. A model with a constant restricted in the cointegration space is estimated for all three cases according to the Johansen (1992) testing methodology.

(*) denotes statistical significance at the 5% critical level.

The roots of the companion matrix

Modulus of 6 largest roots

	USD – DEM					
Unrestricted model	0.95	0.89	0.89	0.80	0.80	0.79
r = 4	1.00	0.92	0.92	0.80	0.80	0.75
	USD – JPY					
Unrestricted model	0.96	0.91	0.91	0.79	0.79	0.76
r = 4	1.00	0.94	0.80	0.78	0.78	0.76

Notes: The table shows the modulus of the estimated p x k roots of the companion matrix from the VAR system, p is the number of variables and k is the number of lags of the VAR. We report the first six roots which are of interest to us.

Table 4: *Tests for overidentifying restrictions*

Case 1:

FED1	FED3	FED12	DEM1 (JPY1)	DEM3 (JPY3)	DEM12 (JPY12)
1	–1	0	0	0	0
1	0	–1	0	0	0
1	0	0	–1	0	0
0	1	0	0	–1	0
0	0	1	0	0	–1

USD – DEM: Q(5) = 0.00; USD – JPY : Q(5) = 0.02

Case 2:

Fed1	Fed3	Fed12	DEM1	DEM3	DEM12
1	–1	0	0	0	0
1	0	–1	0	0	0
*	*	*	*	*	*
*	*	*	*	*	*
*	*	*	*	*	*

USD – DEM: Q(2) = 0.00

Case 3:

Fed1	Fed3	Fed12	DEM1	DEM3	DEM12
*	*	*	*	*	*
*	*	*	*	*	*
1	0	0	–1	0	0
0	1	0	0	–1	0
0	0	1	0	0	–1

USD – DEM: Q(3) = 0.00

Case 4:

Fed1	Fed3	Fed12	DEM1	DEM3	DEM12
0	0	0	1	–1	0
0	0	0	1	0	–1
*	*	*	*	*	*
*	*	*	*	*	*
*	*	*	*	*	*

USD – DEM: Q(2) = 0.03

Notes: Q denotes a likelihood ratio test for overidentifying restrictions as suggested by Johansen and Juselius (1994) and is distributed as a χ^2 with the corresponding degrees of freedom given in parentheses. Numbers in parentheses report marginal significance levels.

Table 5: *Granger – Causality tests*

Variable	JPY3–JPY1		JPY12–JPY1	
DJPY1	0.00	0.00	0.00	0.06
DFED1	0.01	0.05	0.21	0.26

Notes: The null hypothesis is that the variable x does not cause variable y. The first entry in each block refers to the variable x on the corresponding row and y on the corresponding column. For the second entry x and y refer to the variables on the corresponding column and row respectively. The numbers quote the estimated marginal significant level.

Figure 1: *EURO-JPY interest rates for 1, 3 and 12 months, 05/1991-05/2001*

Figure 2: *Fed Funds rates for 1, 3 and 12 months, 05/1991-05/2001*