A Reassessment of the Long-run Validity of the Flexible Price Monetary Exchange Rate Model

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Abstract
In this article we employ the Pesaran and Shin (1999) structural cointegrating VAR methodology to reassess the monetary approach to exchange rate determination. This recently developed technique allows us to test directly the over-identifying restrictions of the long-run structural relations underlying the flexible price monetary model of the exchange rate. Using data for the German Mark-U.S. dollar and the Japanese Yen-U.S. dollar, we find that, for both exchange rates, structural identification is rejected by the data, results that raise further doubts about the long-run validity of the monetary model.

1. Introduction

SINCE the start of the post-Bretton Woods floating era, the monetary exchange rate model has become an influential approach to exchange rate determination. In a sense this is striking in that empirical support for this framework is fraught with controversy. Indeed, though early tests produced results in accordance with theory (for example Hodrick, 1978), later empirical studies (Rasulo and Wilford, 1980; Haynes and Stone, 1981; Driskill and Sheffrin, 1981) found that the monetary approach performed rather poorly. The advent of cointegration analysis highlighted the inadequacy of the standard econometric techniques employed in these studies (which offered spurious inferences due to the non-stationarity of the data employed), and provoked a new spurt of research.

Using the Engle-Granger two-step cointegration procedure, Meese (1986), Baillie and Selover (1987), and McNown and Wallace (1989), among others, were not able to find evidence in support of a long-run equilibrium relationship consistent with the monetary model. Studies which employed the more appropriate Johansen approach to cointegration (MacDonald and Taylor, 1991; Diamandis and Kouretas, 1996; Choudhry and Lawler, 1997; Kouretas, 1997; Makrydakis, 1998), on the other hand, pointed to the existence of one or more cointegrating relations underlying the reduced-form exchange rate equation. In spite of often implausible long-run coefficient estimates leading to the rejection of theory-based restrictions, the latter findings have since been interpreted as evidence in favour of the long-run validity of the monetary model.

We now question this interpretation by arguing that, since in the absence of theory-based restrictions on the underlying structural equations, cointegrating coefficients cannot be identified nor economically interpreted (Wickens, 1996), the 'acid test' for the long-run validity of the monetary model must lie in the structural identification of its cointegrating relations. If the identifying restrictions cannot be rejected through hypothesis testing, then a reduced form exchange rate equation can be obtained from the estimates of the structural equations.

Pesaran and Shin (1999) have provided a full theory of identification that allows testing of theory-based restrictions to obtain, subject to identification, economically interpretable estimates of the long-run structural coefficients in the context of a Johansen-type structural vector autoregressive error correction model (VECM).[2]

In this article we use this recently developed econometric methodology to reassess the long-run validity of the flexible price monetary model for two U.S. dollar bilateral exchange rates, namely, the German Mark and the Japanese Yen. The plan of the paper is as follows. Section 2, reviews the underlying theory, and describes the econometric specification of the flexible price monetary exchange rate model. In Section 3, the structural cointegration VAR methodology is illustrated. The empirical results are presented and discussed in Section 4. Finally, a summary of findings is provided in Section 5.

2. Underlying theory and model specification

The flexible price monetary exchange rate model relies on the continuous purchasing power parity (PPP) hypothesis and an open economy extension of the quantity theory of money. Prices are determined by monetary equilibrium between stable real money demand (which is a function of income and the rate of interest for both economies) and the real money supply. On the assumption that the nominal money stock, income and interest rates are exogenously determined, and that there is perfect mobility of capital and goods, the structural framework is given by the following system of equations:

$$e_t = p_t - p_t^* \qquad (1)$$

$$p_t = \phi_0\, m_t - \phi_1\, y_t + \phi_2\, IR_t \qquad (2)$$

$$p_t^* = \lambda_0\, m_t^* - \lambda_1\, y_t^* + \lambda_2\, IR_t^* \qquad (3)$$

where e_t is the natural logarithm of the nominal spot exchange rate,[3] p_t is the natural logarithm of the price level, m_t is the natural logarithm of the money supply, y_t is the natural logarithm of income, and IR_t is the nominal interest rate. An asterisk denotes a foreign magnitude. Equation (1) is the PPP condition, whereas equations (2) and (3) are the monetary equilibrium conditions for the home and foreign economy, respectively. Solving for the standard reduced-form exchange rate equation we obtain:

$$e_t = \phi_0\, m_t - \lambda_0\, m_t^* - \phi_1\, y_t + \lambda_1\, y_t^* + \phi_2\, IR_t - \lambda_2\, IR_t^* \qquad (4)$$

The model predicts that as the domestic supply of money outstrips the real demand for money in the home economy (due to an increase in the domestic nominal money stock, rising interest rates or lower income levels), monetary equilibrium will be restored through a rise in the domestic price level which will require an increase in the exchange rate (i.e. a depreciation in the price of domestic currency). The literature to date has mainly focussed on testing the validity of the theory's predictions on the reduced form, either on individual cointregrating vectors or across the cointegrating space. The restrictions that are regularly imposed and tested relate to the proportionality between the exchange rate and relative monies (i.e. $\phi_0 = -\lambda_0 = 1$), and equal but opposite signs for the relative income and interest rate coefficients ($\phi_1 = -\lambda_1$ and $\phi_2 = -\lambda_2$, respectively). In this paper we depart from previous studies by testing directly for identification of the underlying structural equations as a pre-requisite for imposing the above restrictions. If the identifying restrictions cannot be rejected then we can solve for an estimated exchange rate equation, and then test the coefficient restrictions across the structural cointegrating vectors rather than on the reduced form model.

All of the data series are obtained from the OECD Main Economic Indicators, except the exchange rate series, which is taken from the Federal Reserve Bank (St. Louis database) and Japanese interest rates which come from the Bank of Japan statistics. Nominal exchange rates are defined as the number of units of domestic currency per unit of US dollar.[4] For prices we use producer price indices, the money supply measure is M1 and industrial production is used as a proxy for income. Interest rates are short-term interest rates. We use seasonally adjusted data for each of the series except the price series.[5] The sample period for estimation is monthly observations from January 1990 to December 1999 for the Japanese Yen/US Dollar exchange rate, and January 1990 to December 1998 for the German Mark/US Dollar. The sample period for the latter ends at 1998 due to the lack of availability of M1 data for Germany after this period.

3. Econometric methodology

Until recently it has proved difficult to give an economic interpretation to multiple cointegrating vectors. This is because they are obtained from the estimation of a reduced-form system of variables using a Vector Autoregressive Error Correction Model (VECM). The size of the cointegrating space is defined by the number of structural equations underlying the reduced form, with one cointegrating vector for each endogenous variable. Wickens (1996) proved that in the absence of identifying restrictions, the estimated cointegrating vectors will not be identified, nor will the associated common stochastic trends. When r, the number of cointegrating vectors, equals unity, the 'normalizing' or 'just-identifying' restriction (suggested originally from the work of Johansen 1991, 1995) will be sufficient to identify the long-run model from any of the integrated variables in the system. This is because the reduced form is the same as the structural equation. Pesaran and Shin (1999) show that for each cointegrating vector there should be the same number of restrictions as the size of the cointegrating space (i.e. the total number of cointegrating vectors). Thus the minimum condition for exact-identification is that the total number of restrictions, k should equal r^2. In this application, economic theory suggests that three cointegrating vectors should exist from the three structural equations. Consequently, nine restrictions are needed for exact-identification.[6] Pesaran and Shin (1999) suggest that the r^2 restrictions will consist of r 'normalizing' restrictions and r^2-r additional restrictions derived from economic theory. The exact-identifying hypothesis provides a test for statistical identification when there is more than one cointegrating vector. However, by imposing all of the remaining restrictions from economic theory, the over-identifying hypothesis can be used to establish whether a full structural representation can be obtained from the cointegrating relations. We show that in this application a total of twenty three restrictions need to be imposed, fourteen of which are used to test the over-identifying hypothesis.

To consider the problem of structural identification, the system in (1) to (3) is re-parameterised as p-order VAR in levels:

$$B_0 z_t + B_1 z_{t-1} + B_2 z_{t-2} + \dots + B_p z_{t-p} + \phi D_t = \varepsilon_t \quad (5)$$

Where z_t $[z_t = w_t', x_t']'$ is a $(m+n)\times 1$ vector of $I(1)$ variables, which consists of an $m\times 1$ vector of endogenous variables, w_t $[w_t = (e_t, p_t, p_t^*)']$ and an $n\times 1$ vector of weakly exogenous variables, x_t $[x_t = (m_t, m_t^*, y_t, y^*, IR_t, IR)']$. D_t is a vector of exogenous variables.[7] For cointegration analysis, (5) is re-expressed as a structural VECM:

$$A_0 \Delta z_t = \sum_{i=1}^{p-1} A_i \Delta z_{t-i} + \pi^* z_{t-p} + \phi D_t + u_t \quad (6)$$

where A_i is an (m+n)xp matrix of coefficients for the underlying VAR of Δz_t to Δz_{t-i}, where $A_i = -\sum_{j=i+1}^{p} B_j$ and π^* is the long-run matrix of structural coefficients, defined as $\pi^* = A_0 - \sum_{i=1}^{p} A_i$. Provided $r \geq 1$ cointegrating vectors exist, then we can factorise π^* as $\pi^* = \alpha^* \beta^*$, where α^* is a matrix of adjustment coefficients and β^* is the long-run coefficient matrix for the cointegrating vectors. Estimation of the cointegrating vectors by the Johansen procedure involves re-expressing (6) in a reduced form:

$$\Delta z_t = \sum_{i=1}^{p-1} \Gamma_i \Delta z_{t-i} + \pi z_{t-p} + \Phi D_t + \upsilon_t \qquad (7)$$

where $\Gamma_i = A_i A_0^{-1}$, $\pi = \pi^* A_0^{-1}$, $\Phi = \phi A_0^{-1}$, and $\nu_t = u_t A_0^{-1}$. The number of distinct cointegrating vectors can be found by testing for the rank of π using the conventional likelihood ratio testing procedure.

Given we found that some of the variables used were trended we opted to estimate the VECM with unrestricted intercepts and restricted trends:

$$\Delta z_t = \alpha_0 + \sum_{i=1}^{p-1} \Gamma_i \Delta z_{t-i} + \alpha\beta' (Z_{t-p} - \gamma T) + \Phi D_t + \nu_t \quad (8)$$

where α_0 is a $(m+n)\times 1$ matrix of drift coefficients, T are time trends and γ is an $r\times 1$ vector of constants. By estimating the VECM in this form we allow for the possibility of linear trends in the cointegrating relations. The vector of constants (γ) are used to restrict the β coefficients for the trend terms and thus avoid the possibility that when the model is expressed in levels of z_t the nature of the trends will vary with the number of cointegrating vectors (Pesaran and Shin, 1999). The estimated coefficients from the VECM are obtained by maximum likelihood estimation.

To obtain the long-run structural coefficient matrix (β^*) at least r^2 exactly-identifying restrictions need to be imposed on the β matrix from (8). Thus the necessary and sufficient condition for exact-identification is:

$$\text{Rank}\ (R_i\, \beta_i) = r \qquad (9)$$

Where R_i is the matrix of restrictions for the ith cointegrating vector, βi. There must be r restrictions for each cointegrating vector. Thus from (1) to (3) a total of nine restrictions need to be imposed, three for each cointegrating vector.

To consider the structural representation of the vector z_t and the time trend T it is convenient to use the following representation:

$$\beta_{11}\, e_t = \alpha_{10} + \beta_{21}\, p_t + \beta_{31} + \beta_{101}\, T + \xi_{1t} \quad (10)$$

$$\beta_{22}\, p_t = \alpha_{20} + \beta_{42}\, m_t + \beta_{62}\, y_t + \beta_{82}\, IR_t + \beta_{102}\, T + \xi_{2t} \quad (11)$$

$$\beta_{33}\, p_t^* = \alpha_{30} + \beta_{53}\, m_t^* + \beta_{73}\, y_t^* + \beta_{93}\, IR_t^* + \beta_{103}\, T + \xi_{3t} \quad (12)$$

where ξ_{1t}, ξ_{2t}, ξ_{3t} are random errors for each structural equation. (10) to (12) are consistent with the matrix of long-run cointegrating vectors:

$$\beta = \qquad (13)$$

$$\begin{pmatrix} \beta_{11} & \beta_{21} & \beta_{31} & \beta_{41} & \beta_{51} & \beta_{61} & \beta_{71} & \beta_{81} & \beta_{91} & \beta_{101} \\ \beta_{12} & \beta_{22} & \beta_{32} & \beta_{42} & \beta_{52} & \beta_{62} & \beta_{72} & \beta_{82} & \beta_{92} & \beta_{102} \\ \beta_{13} & \beta_{23} & \beta_{33} & \beta_{43} & \beta_{53} & \beta_{63} & \beta_{73} & \beta_{83} & \beta_{93} & \beta_{103} \end{pmatrix}$$

corresponding to the vector

$$z_t = (e_t, p_t, p_t^*, m_t, m_t^*, y_t, y_t^*, IR_t, IR_t^*)'$$

and the time trend T.

One null hypothesis for exact-identification can be given as:

$$H_E: \begin{Bmatrix} \beta_{11} = -1 & \beta_{41} = 0 & \beta_{51} = 0 \\ \beta_{12} = 0 & \beta_{22} = -1 & \beta_{52} = 0 \\ \beta_{13} = 0 & \beta_{23} = 0 & \beta_{33} = -1 \end{Bmatrix} \qquad (14)$$

The exactly-identifying hypothesis can be explained as follows. We use the first cointegrating vector to test for the PPP condition, while the second and third vectors are used to identify the domestic and foreign money balances equations. For the first cointegrating vector β_{11} = -1 is the 'normalizing' restriction for the exchange rate equation, while β_{41} = 0 and β_{51} = 0 since m_t and m_t^* do not appear in the PPP condition. For the second cointegrating vector, β_{22} = -1 is the 'normalizing' restriction for p_t, while $\beta_{12} = 0$ and $\beta_{52} = 0$ since e_t and m_t^* do not appear in the domestic price equation. Finally, $\beta_{33} = -1$ is the normalizing restriction for p_t^* and $\beta_{13} = 0$ and $\beta_{23} = 0$ since e_t and p_t do not appear in the structural equation for p_t^*. The exactly-identifying hypothesis is tested by comparing the maximized value of the log-likelihood function for the VECM under the exactly-identifying restrictions with the maximized log-likelihood from the unrestricted VECM. However, it should be noted that the exactly-identifying hypothesis is not unique and could be obtained from imposing any nine of the restrictions implied by economic theory. Moreover, it only provides a statistical test for identification. To test whether the estimated cointegrating vectors are fully consistent with the structural representation implied by economic theory, an additional d over-identifying restrictions need to be imposed. Thus the total number of restrictions is $k=d+r^2$. Specifically for the first cointegrating vector exclusion restrictions need to be imposed on the coefficients for y_t, y_t^*, IR_t and IR_t^* and the restrictions $\beta_{21} = 1$ and $\beta_{31} = -1$ is imposed to satisfy the PPP condition. In the second cointegrating vector, exclusion restrictions are imposed on the three coefficients of p_t^*, y_t^* and IR_t^* and a unity restriction for the coefficient of m_t. Finally, the third cointegrating vector has a zero restriction for the coefficients of m_t, y_t, and IR_t and $\beta_{53} = 1$ for the m_t^*, variable. Thus the null hypothesis for over-identification is:

$$H_0 = \begin{pmatrix} -1 & 1 & -1 & 0 & 0 & 0 & 0 & 0 & 0 & \beta_{101} \\ 0 & -1 & 0 & 1 & 0 & \beta_{62} & 0 & \beta_{82} & 0 & \beta_{102} \\ 0 & 0 & -1 & 0 & 1 & 0 & \beta_{73} & 0 & \beta_{93} & \beta_{103} \end{pmatrix} \quad (15)$$

Structural identification and estimation of the cointegrating vectors is achieved through maximisation of a log-likelihood function subject to the exactly- and over-identifying restrictions. The null hypothesis is tested using the following L-R test statistic:

$$\text{L-R} = 2\,(LR_E - LR_O) \sim \chi_r^2(d) \quad (16)$$

where LR_E is the maximised value of the log-likelihood function under the exactly-identifying restrictions and LR_O is the maximised value of the log-likelihood function under both exactly- and over-identifying restrictions. This L-R test statistic has an asymptotically valid chi-squared distribution with $d=k-r^2$ degrees of freedom. If we cannot reject the null hypothesis a structural interpretation of the estimated coefficients of the cointegrating vectors is ensured.

4. Estimation results

To ensure that all of the variables used are $I(1)$, Augmented Dickey-Fuller (ADF) (Dickey and Fuller, 1979) and Phillips-Perron (PP) (Phillips and Perron, 1988) tests were calculated. The tests suggest that the level of each variable is $I(1)$ and first difference stationary (see Table 1). Some of the series were subject to one-period shocks and in these cases the Perron (1989) unit root test was used. Again the level of each variable was found to be $I(1)$.

In order to define the size of the cointegrating space, an unrestricted VECM was estimated with unrestricted intercepts and restricted trends. Five tests of the cointegrating rank were employed: the trace and maximum eigenvalue statistics, and the AIC, SBC and HQC (Hannan and Quinn, 1979) selection criteria. The estimation results are shown in Table 3. For both exchange rates the null hypothesis of $r=0$ (no cointegration) is rejected. However, the selection procedures do not provide complete agreement as to the exact size of the cointegrating space. For both exchange rates we can draw the following conclusions: (i) $r=5$ from the maximum eigenvalue statistic and HQC; (ii) $r=4$ from the trace statistic and SBC. The AIC suggests $r=8$ for the German Mark/US Dollar and $r=5$ for the Japanese Yen/US Dollar. These results fail to provide empirical support for the long-run theory prediction, namely, that there should be three cointegrating relations: one to represent the

Table 1: Unit Root Tests

DM/$US Exchange Rate (1990m1-1998m12)

Variable	*Augmented Dickey-Fuller tests*		*Phillips-Perron tests*	
	Level	First difference	Level	First difference
e	-2.5100 (1)	-6.8362 (0)	-1.4451	-7.7326
p^*	-2.5001 (0)	-8.7721 (0)	-1.8046	-8.4328
m^*	-0.8538 (1)$^{\tau}$	-3.9590 (0)	-0.4903$^{\tau}$	-4.5909$^{\tau}$
y^*	-3.1079 (0)$^{\tau}$	-8.6226 (0)	-2.8558$^{\tau}$	-8.8591$^{\tau}$
IR^*	-2.0775 (1)	-6.1358 (0)	-2.0388	-6.2349$^{\tau}$
p	-2.4852 (3)	-8.6039 (0)	-2.8417	-8.7586
m	-0.8951 (0)¢	-17.129 (0)¢	-	-
y	-1.4863 (0)	-13.1318 (0)	-0.7649	-14.747
IR	-1.0282 (1)	-7.6223 (0)	-0.3178	-10.863

Yen/$US Exchange Rate (1990m1-1999m12)

Variable	*Augmented Dickey-Fuller tests*		*Phillips-Perron tests*	
	Level	First difference	Level	First difference
e	-1.9396 (1)	-7.2947 (0)	-1.6545	-7.2789
p^*	-1.9202 (1)	-9.1369 (0)	-1.1728	-8.7780
m^*	-1.3954$^{\tau}$ (1)	-4.1435 (0)	-1.2311$^{\tau}$	-3.9533
y^*	-3.3620$^{\tau}$ (0)	-9.1229 (0)	-3.0817$^{\tau}$	-8.8839
IR^*	-2.1285 (1)	-6.8167 (0)	-2.0489	-6.6709
p	-4.5271 (2)†	-5.8673 (0)†	-	-
m	-1.3554 (0)§	-6.1984 (0)§	-	-
y	-1.7238 (3)	-6.2014 (2)	-2.2440	-17.281
IR	-1.1525 (3)	-4.7596 (2)$^{\mu}$	-	-

Critical values

Significance level	*Drift component*	*Drift and trend components*
90%	-2.57	-3.13
95%	-2.89	-3.45

Notes: In each case the USA is the foreign economy and Germany/Japan the domestic economy. τ denotes a significant trend component. The number of lagged terms used in the ADF regression are shown in parentheses. The truncation lag for the Phillips-Perron tests is set from the highest significant lag term in the autocorrelation function for each variable.

¢ denotes that the German M1 variable (*m* in the first table) was subject to a structural break at 90M6 and 90M12. The 90% and 95% critical values for the Perron test are -3.40 and -3.86 respectively.

† denotes that the Japanese price variable (p in the second table) was subject to a structural break at 97M4; § indicates Japanese M1 (*m*) had a break at 90M5; while μ indicates a break to Japanese interest rates (*IR*) between 93m3 and 94m4. The 90% and 95% Perron (1989) critical values for Japanese prices are -5.32 and -3.87; for Japanese M1 -4.57 and -3.40; and Japanese interest rates -4.28 and -2.83.

Table 2: Lag selection criteria

DM/$US (1990m1-1998m12)					Yen/$US (1990m1-1999m12)				
Var order (p)	*Log likelihood*	*AIC*	*SBC*	Notes: AIC and *Adjusted L-R test*	*Var order (p)*	*Log likelihood*	*AIC*	*SBC*	*Adjusted L-R test*
10	4508.6	3662.6	2528.1	--	10	4403.4	3539.4	2335.2	--
9	4039.4	3274.4	2248.4	$\chi^2_{81}=121.66$	9	4158.3	3375.3	2284.0	$\chi^2_{81}=98.04$
8	3793.8	3109.8	2192.5	$\chi^2_{162}=185.34$	8	3996.3	3294.3	2315.9	$\chi^2_{162}=162.84$
7	3614.7	3011.7	2203.0	$\chi^2_{243}=231.77$	7	3865.8	3244.8	2379.3	$\chi^2_{243}=215.03$
6	3469.5	2947.5	2247.4	$\chi^2_{324}=269.41$	6	3765.6	3225.6	2472.9	$\chi^2_{324}=255.13$
5	3342.8	2901.8	2310.4	$\chi^2_{405}=302.24$	5	3672.1	3213.1	2573.4	$\chi^2_{405}=292.52$
4	3253.5	2893.5	2410.7	$\chi^2_{486}=325.40$	4	3608.2	3230.2	2703.4	$\chi^2_{486}=318.06$
3	3182.5	2903.5	2529.3	$\chi^2_{567}=343.81$	3	3496.7	3199.7	2785.7	$\chi^2_{567}=362.68$
2	3094.6	2896.6	2631.1	$\chi^2_{648}=366.60$	2	3423.2	3207.2	2906.1	$\chi^2_{648}=392.09$
1	3006.0	2889.0	2732.0	$\chi^2_{729}=389.58$	1	3327.9	3192.9	3004.7	$\chi^2_{729}=430.21$
0	1737.5	1701.5	1653.2	$\chi^2_{810}=718.45$	0	1965.9	1911.9	1836.7	$\chi^2_{810}=974.98$

Notes: AIC and SBC are the Akaike Information and Schwarz Bayesian lag selection criteria. The likelihood ratio test statistic is adjusted for degrees of freedom, which tests the null hypothesis that the VAR order is p (p=0,.....,4) against the alternative of p+1. Under the null the test statistic has an asymptotic chi-squared distribution with $n^2(p+1-p)$ degrees of freedom, where n is the number of variables in the system.

PPP condition and one each for the domestic and foreign money balance equations. However, in view of the contradictory evidence and because cointegration tests may over-reject the null hypothesis in finite samples, we decided to proceed assuming r=3, as a further test of the economic theory.

The coefficients of the cointegrating vectors from the unrestricted VECM are presented in Table 4. Clearly, without identifying restrictions, it is difficult to interpret the sign and magnitude of each coefficient in accordance with economic theory. Moreover, no individual cointegrating vector can be used to represent the monetary exchange rate equation, without first identifying the structural system and then solving for the estimated exchange rate equation.

We therefore proceeded to impose the nine exactly-identifying restrictions defined by the null hypothesis in (14), from which the maximized log-likelihood was obtained. The estimates for the German Mark/US Dollar and Japanese Yen/US Dollar are 2896.2 and 3221.8. Both of these values are identical to the maximized log-likelihood from the unrestricted VECMs (see Table 2). Thus we cannot reject the null hypothesis of exact-identification.

The over-identifying hypothesis (15) was tested by imposing an additional 14 over-identifying restrictions on the cointegrating space as suggested by the theory underpinning (1) to (3). The maximum likelihood estimates of the structural cointegrating vectors are shown in Table 5. The estimated likelihood ratio test statistic for the over-identifying restrictions is 81.25 for the German Mark/US Dollar, and 173.47 for the Japanese Yen/US Dollar. Since both of the estimates exceed the 95 per cent chi-squared critical value for 14 degrees of freedom (23.68), the identification testing procedure indicates that the structural restrictions implied by the monetary model of the exchange rate are rejected by the data.

Statistical rejection of structural identifica-

Table 3: Cointegration Rank Test Statistics

Null	*Alternative*	*Maximum Eigenvalue Statistic*				*Trace Statistics*			
H_0	H_1	*DM/\$US*	*Yen/\$US*	*95% c.v.*	*90% c.v.*	*DM/\$US*	*Yen/\$US*	*95% c.v.*	*90% c.v.*
r=0	r=1	142.62	139.69	61.27	58.09	442.83	455.70	222.62	215.87
r≤1	r=2	80.69	103.89	55.14	52.08	300.22	316.00	182.99	176.92
r≤2	r=3	78.02	83.19	49.32	46.54	219.53	212.11	147.27	141.82
r≤3	r=4	59.59	58.74	43.61	40.76	141.51	128.92	115.85	110.60
r≤4	r=5	43.59	36.05	37.86	35.04	81.91	70.18	87.17	82.88
r≤5	r=6	16.00	13.89	31.79	29.13	38.31	34.13	63.00	59.16
r≤6	r=7	12.99	9.39	25.42	23.10	22.31	20.24	42.34	39.34
r≤7	r=8	8.24	7.33	19.22	17.18	9.31	10.85	25.77	23.08
r≤8	r=9	1.08	3.52	12.39	10.55	1.08	3.52	12.39	10.55

Model Selection Criteria

	DM/\$US				*Yen/\$US*			
rank	*maximised log-likelihood*	*AIC*	*SBC*	*HQC*	*maximised log-likelihood*	*AIC*	*SBC*	*HQC*
r=0	2784.5	2757.5	2721.3	2742.9	3100.0	3055.0	2992.3	3029.6
r=1	2855.8	2810.8	2750.5	2786.4	3169.9	3106.9	3019.1	3071.2
r=2	2896.2	2835.2	2753.5	2802.0	3221.8	3142.8	3032.7	3098.1
r=3	2935.2	2860.2	2759.6	2819.4	3263.4	3170.4	3040.8	3117.8
r=4	2965.0	2878.0	2761.3	2830.7	3292.8	3187.8	3041.4	3128.4
r=5	2986.8	2889.8	2759.7	2837.0	3310.8	3195.8	3035.5	3130.7
r=6	2994.8	2889.8	2749.0	2832.7	3317.8	3194.8	3023.3	3125.1
r=7	3001.3	2890.3	2741.4	2829.9	3322.5	3193.5	3013.7	3120.4
r=8	3005.4	2890.4	2736.2	2827.9	3326.1	3193.1	3007.7	3117.8
r=9	3006.0	2889.0	2732.0	2825.3	3327.9	3192.9	3004.7	3116.5

Note: The critical values are asymptotic values obtained from Pesaran, Shin and Smith (2000).

Table 4: Unrestricted Cointegrating Vectors

e_t	p_t	p_t^*	m_t	m_t^*	y_t	y_t^*	IR_t	IR_t^*	T
					DM/\$US				
-0.482	0.963	-4.503	0.387	-1.181	-0.433	-0.613	0.019	-0.046	0.0064
-0.274	-1.273	-2.995	-0.911	-0.269	-0.542	-0.216	-0.031	-0.081	0.0107
-0.799	1.751	-5.520	0.106	-3.065	-3.759	-6.427	0.017	0.053	0.038
					Yen/\$US				
-0.559	3.433	-3.008	0.1129	-1.309	0.0029	0.5441	-0.00015	-0.0629	0.0039
0.152	1.801	-3.746	1.047	1.774	1.476	-0.093	0.017	-0.038	-0.0025
0.114	2.376	-3.631	-2.773	-4.605	-0.987	-9.551	-0.197	0.105	0.053

Table 5: Cointegrating Vectors subject to exact- and over-identifying restrictions

DM/$US

e_t	p_t	p_t^*	m_t	m_t^*	y_t	y_t^*	IR_t	IR_t^*	T
-1	1	-1	0	0	0	0	0	0	0.000539
0	-1	0	1	0	1.5553 (2.8603)	0	0.3479 (0.4635)	0	0.005732 (0.0174)
0	0	-1	0	1	0	3.1649 (1.2257)	0	-0.0547 (0.0341)	-0.01279 (0.0045)

Likelihood ratio test of over-identifying restrictions: $\chi^2_{14} = 81.25$

Yen/$US

e_t	p_t	p_t^*	m_t	m_t^*	y_t	y_t^*	IR_t	IR_t^*	T
-1	1	-1	0	0	0	0	0	0	0.0025 (0.0019)
0	-1	0	1	0	-1.2859 (0.7308)	0	0.0471 (0.0330)	0	-0.0062 (0.0012)
0	0	-1	0	1	0	1.9673 (0.6172)	0	-0.0156 (0.0079)	-0.00632 (0.0022)

Likelihood ratio test of over-identifying restrictions: $\chi^2_{14} = 173.47$

Estimated standard errors in parentheses

tion means that no economically meaningful interpretation can be given to the estimated coefficients (which, not surprisingly, appear to be incorrectly signed). The estimation results, therefore, suggest that while more than one long-run relation exists among the variables suggested by theory, the nature of the underlying structural system is not consistent with that postulated by the flexible price monetary exchange rate model. These findings go some way to explain why the theory-based, reduced-form restrictions imposed on cointegrating vectors in earlier studies have in the main been rejected through hypothesis testing.

5. Conclusions

In this article, we reassessed the validity of the monetary approach to exchange rate determination by testing for identification of the long-run structural relations underlying the flexible price exchange rate model. Using data for the German Mark and Japanese Yen bilateral US dollar rates, we found that, for both exchange rates, the over-identifying restrictions were rejected by the data, results which suggest that the flexible price monetary model is not a valid framework for analysing the long-run exchange rate.

Endnotes

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2. For recent applications of this methodology see Abbott and De Vita (2001), Garratt *et al.* (1999, 2000), Pesaran, Shin and Smith (2000).

3. Defined as the number of units of domestic currency per US dollar.

4. Exchange rates are monthly averages of daily observations.

5. To test for seasonality, each price series was regressed on a constant and eleven seasonal dummies. An F-test was used to test for the significance of the dummy coefficients and hence to detect an overall seasonal pattern in the data. We found no evidence of any seasonal movements in any of the price series used.

6. We can see how general this result is, since if there is a unique cointegrating vector then only one restriction needs to be imposed, which is the Johansen 'normalizing' restriction.

7. In our application, the exogenous components take the form of dummy variables, which are used to encompass structural breaks to some of the series used, when estimating the VECM.

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