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IS THE EXCHANGE RATE PREDICTABLE IN THE LONG-RUN?

SOOWON MO*and WOOKIL CHO**

The short-run and long-run forecasts of exchange rates based on the three structural models are compared to those based on the random walk model. The long-run forecasts are generated by the error-correction equations of the Johansen's multivariate cointegration technique. The results show that while the random walk model outperforms the structural models in the short-run forecasting, the structural models outperform the random walk in the long-run forecasting. Our results indicate that the monetary models should be thought more of as the long-run models.

1. INTRODUCTION

This paper evaluates the forecasting accuracy of the monetary models of exchange rate determinaiton. The monetary approach to exchange rate determination views the current exchange rate as being determined by the interaction of the outstanding stocks of and demand for the two monies. There are three variants of the monetary models of exchange rates. The first is the "flexible-price" model of Frenkel(1976) and Bilson (1978) which relies on the twin assumptions of purchasing power parity and the existence of stable money demand functions for the domestic and foreign economies. The second is the "sticky-price" model developed by Dornbusch (1976) and extended by Frankel (1979) to allow for sustained inflation differentials across countries. The third variant is the Hooper-Morton's (1982) sticky-price model that allows for shifts in the long-run equilibrium real exchange rate and explains the relationship between the real exchange rate and the current account.

The Frenkel-Bilson, Dornbusch-Frankel, and Hooper-Morton forms of the monetary model have been the subjects of empirical tests of Meese and Rogoff (1983). Meese and Rogoff compared the forecasting performance of the structural models with that of the random walk for dollar/DM, dollar/Pound, dollar

* Associate Professor of International Trade, Mokpo National University.

** Associate Professor of International Trade, Sooncheon National University.

/Yen, and the trade−weighted dollar exchange rates, using the data for the period March 1973 through June 1981. Their result casts serious doubt on the ability of the monetary models to predict exchange rate movements.

In this paper, we study the forecasting ability of the monetary models from a different perspective. We investigate the forecasting performance of the monetary models in the short run and in the long run, and show that while the structural monetary models provide poorer performance than the random walk for short−run forecasting, they outperform the random walk model in the long −run forecasting. It is shown by using an multivariate cointegration technique proposed by Johansen (1988) and Johansen and Juselius (1990).

The remainder of the paper is organized as follows: Section 2 briefly presents the Johansen−Juselius procedure of multivariate cointegraiton. Section 3 specifies the monetary models which are employed in the empirical work; Section 4 reports the econometric evidence. In section 5 the paper provides the concluding remarks.

2. METHODOLOGY

The long−run relationship between exchange rate and the economic variables can be investigated by the two−step regression technique of Engle and Granger (1987). One problem with the EG test is that the cointegrating relationship may not be unique when there are more than two variables in the system. A relatively new test is made also for cointegration, using the maximum likelihood methodology of Johansen (1988) and Johansen and Juselius (1990). The particular advantages of this approach are that it captures fully the underlying time−series properties of the data and also offers a test statistic which has an exact limiting distribution. This methodology is briefly outlined here. Consider the following vector autoregressive model of order k:

$$(1)\ X_t = \Pi_1 X_{t-1} + \cdots + \Pi_k X_{t-k} + \mu + \varepsilon_t,\ (t = 1, \cdots, T)$$

where μ denotes drift and $\varepsilon_1, \cdots, \varepsilon_T$ are $IIN_p(0, \Lambda)$. In general, economic time series are non−stationary processes, and VAR systems have usually been expressed in first differenced form.

$$(2)\ \Delta X_t = \Gamma_1 \Delta X_{t-1} + \cdots + \Gamma_k \Delta X_{t-k+1} + \Pi X_{t-k} + \mu + \varepsilon_t$$

where

$$\Gamma_i = -(I - \Pi_1 - \cdots - \Pi_i),\ (i = 1, \cdots, k-1)$$

$$(3)\ \Pi = -(I - \Pi_1 - \cdots - \Pi_k).$$

The long−run relationship between the variables in the data vector is contained in the coefficient matrix Π. If the matrix Π is the null matrix, i.e. rank $(\Pi)=0$, there is no long-run relationship among the variables and (2) corresponds to a traditional differenced vector time series model.[1] If the matrix Π is of full rank, it indicates that the vector process X_t is stationary. We can write $\Pi = \alpha\beta'$ where β is a matrix of cointegrating vectors and α is a matrix of error correction coefficients if $0<\text{rank}(\Pi)=r<p$, where p is the number of variables in the system.

The likelihood ratio test statistic, the trace test, for the hypothesis that there are at most r cointegrating vectors against the general unrestricted model r=p is

$$\text{trace}(r) = -2\ln(Q) = -T\sum_{i=r+1}^{p}\ln(1-\tilde{\lambda}_i)$$

where $\tilde{\lambda}_i$ correspond to the p−r smallest squared canonical correlations between the residuals of X_{t-k} and ΔX_t. The p×r matrix of the cointegrating vectors β can be obtained as the p−element eigenvectors correponding to the r largest eigenvalues.

The LR statistic for testing that there are r versus r+1 cointegrating vectors, the maximal eigenvalue test, is given by

$$\lambda_{max}(r) = -2\ln(Q\,;\,r\mid r+1) = -T\ln(1-\tilde{\lambda}_{r+1})$$

Osterwald−Lenum (1992) tabulated the distribution of trace(r) and $\lambda_{max}(r)$.[2]

3. THE MONETARY MODELS

As representative structural models Meese and Rogoff choose the flexible−price monetary (Frenkel−Bilson) model, the sticky−price monetary (Dornbusch−Frankel) model, and the stcky−price asset (Hooper−Morton) model. One general specification covering all these models can be written as follows:

$$(4)\ s_t = a_0 + a_1 m_t + a_2 m_t^* + a_3 y_t + a_4 y_t^* + a_5 i_{st} + a_6 i_{st}^* + a_7 i_{Lt} + a_8 i_{Lt}^* + a_9 CTB_t + a_{10} CTB_t^*$$

where s = the log of the spot exchange rate, m = log of the domestic money supply, y = log of real income, i_s = short−term interest rate, i_L = long−term interest rate, and CTB = cumulated trade balance. The asterisk denotes the correponding foreign variable.

The Frenkel−Bilson model assumes $a_7 = a_8 = 0$. The Dornbusch−Frankel

1) Notice that model(2) is expressed as a traditional first difference VAR model except for the term ΠX_t

2) The tables in Johansen(1988) and Johansen and Juselius(1990) have been recalculated and extended to handle a full test sequence from full rank to zero rank for at most 11−dimensional systems($p\leq 11$). In addition two new cases(2* and 2) have been tabulated.

model, which allows for slow price adjustment, constraints $a_9 = a_{10} = 0$.
The conventional and parsimonious assumption is to impose the constraints that the domestic and foreign money demand coefficients are equal. The demand for money, which is the central behavioral equation of the monetary approach, is given by

$$m_t = p_t + \Phi y_t - \lambda i_t$$

$$m_t^* = p_t^* + \Phi^* y_t^* - \lambda^* i_t^*$$

where

$\Phi(\Phi^*) \equiv$ the money demand elasticity with respect to income,

$\lambda(\lambda^*) \equiv$ the money demand semielasicity with respect to the interest rate.

If the restictions that $\Phi = \Phi^*$ and $\lambda^* = \lambda^*$ are not met, then the domestic variable and its correponding foreign variable should enter the monetary equation in the separated form.

Table 1 gives SUR (Seemingly Unrelated Regression) estimates of the money demand function. The data used in the money demand function consist of monthly observations of M1 (M2 for U.K.), the call money rate (prime corporate paper rate for Canada), and seasonally adjusted industrial production. The monthly data cover the period April 1973 through June 1993. All data are taken from the OECD Main Economic Indicators. The test statistics strongly reject the hypothesis that $\Phi = \Phi^*$ and $\lambda^* = \lambda^*$. This leads to a specification of the model that separates the domestic and foreign variables.

Table 1. Statistics of demand for money

restriction	Canada	France	Germany	Japan	U.K.
$\Phi = \Phi^*$	39.13	38.00	472.9	1056	139.9
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
$\lambda = \lambda^*$	16.21	33.14	42.01	227.8	62.11
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)

Notes: 1. The F test for equality of coefficients of each country is aginst the U.S.
2. Marginal significance levels are reported in parentheses below each F statistics.

This paper also replaces the cumulated trade balance of Hooper and Morton's specification with the current trade balance. It has three reasons. The first is that the consistent estimation of the cointegrating vector via OLS is possible for only the case where all variables are I(1). The second is that the tests employed in this paper are only valid if the variables have the same order of integration (Stock 1987). And the third reason can be found in Table 2, which reports the cross correlations of current trade balance and cumulated trade balance estimated from 0 lag to 18 lags. These calculations reveal that trade balance and cumulated trade balance are, by and large, highly correlated.

Table 2. Cross−correlations of current trade balance and cumulated trade balance

country	0	-1	-2	-3	-4	-6	-8	-10	-12	-14	-16	-18
Canada	0.62	0.62	0.62	0.62	0.63	0.64	0.65	0.67	0.69	0.70	0.71	0.72
France	0.53	0.54	0.56	0.58	0.60	0.60	0.61	0.63	0.62	0.62	0.61	0.67
Germany	0.63	0.64	0.66	0.67	0.68	0.70	0.72	0.76	0.80	0.81	0.83	0.83
Japan	0.77	0.76	0.75	0.75	0.74	0.73	0.71	0.69	0.69	0.69	0.67	0.69
U.K.	0.69	0.70	0.71	0.71	0.71	0.72	0.73	0.73	0.73	0.73	0.74	0.74
U.S.	0.71	0.72	0.72	0.72	0.73	0.75	0.76	0.76	0.77	0.78	0.81	0.82

Before implementing tests of cointegration, we proceed by examining the time series properties. Since all variables are time series, the first step is to determine if the variables are stationary in levels, or if a first differencing is required to achieve stationarity. The results on the nonstationarity of the data from the application of the augumented Dickey−Fuller (ADF) test and Phillips−Perron test are presented in Table 3.

The Phillips−Perron test, which is a modified DF test, is based on a non−parametric correction for any possible serial correlation or time−dependent heteroscedasticity in the residual. The Phillips−Perron test statistic is given by

$$Z_t = (s_u/s_{TL})t\hat{\beta} - (s^2_{TL} - s^2_u)/2\left[T^{-1}s_{TL}\left(\sum_{t=1}^{T} X^2_{t-1}\right)^{1/2}\right]^{-1}$$

where $t\hat{\beta}$ is the usual t−statistic in the simple Dickey−Fuller (DF) test; s^2_u is the sample variance of the estimated residuals; s^2_{TL} is a consistent variance estimator given by

$$s^2_u = T^{-1}\sum_{t=1}^{T} u^2_t$$

$$s^2_{TL} = T^{-1}\sum_{t=1}^{T} u^2_t + 2T^{-1}\sum_{k=1}^{L} w_{kL}\sum_{t=k+1}^{L} u_t u_{t-k}$$

where L is called the lag truncation number and $w_{KL} = 1 - k/(L+1)$ denotes the weighting scheme. Newey and West (1987) have suggested a modification to variance estimators such as s^2_{TL} which ensures nonnegative variance estimate. It is because s^2_{TL} can take on negative values when there are large negative sample serial covariances (Phillips 1987).

As mentioned above, the Phillips–Perron test is robust to a wide variety of serial autocorrelation and time–dependent heteroskedasticity. Schwert(1987) suggests, however, on the basis of Monte–Carlo evidence, that the Phillips–Perron tests may be biased toward rejecting the null hypothesis of non–stationarity much too frequently[3]. As a cross–check, therefore, ADF test is necessary to employ. To test for a unit root in the series Xt the following regressions are estimated via OLS.

$$DX_t = \beta X_{t-1} + \sum_{j=1}^{p} \beta_j DX_{t-j} + \varepsilon_t$$

$$DX_t = \beta_0 + \beta X_{t-1} + \sum_{j=1}^{p} \beta_j DX_{t-j} + \varepsilon_t$$

$$DX_t = \beta_0 + \beta X_{t-1} + \sum_{j=1}^{p} \beta_j DX_{t-j} + \mu_t t + \varepsilon_t$$

To test for a unit root in the series DXt the following regressions are also estimated via OLS.

$$D^2X_t = \beta DX_{t-1} + \sum_{j=1}^{p} \beta_j D^2X_{t-j} + \varepsilon_t$$

$$D^2X_t = \beta_0 + \beta DX_{t-1} + \sum_{j=1}^{p} \beta_j D^2X_{t-j} + \varepsilon_t$$

$$D^2X_t = \beta_0 + \beta DX_{t-1} + \sum_{j=1}^{p} \beta_j D^2X_{t-j} + \mu_t t + \varepsilon_t$$

where D is the first difference operator (i.e., $DX_t = X_t - X_{t-1}$) and D^2X_t denotes $DX_t - DX_{t-1}$. The number of lags entering the estimated equations is determined on the Ljung–Box Q–statistic for serial correlation. Starting with the longest lags ($p=15$), augumentation terms are eliminated from the specification unless their elimination introduces serial correlation.

The results of both ADF and Phillips–Perron tests indicate that the null hypothesis of the unit root cannot be rejected for the (log)level variables with a few exceptions. The ADF $t\alpha^{\sim}$ statistic of Japanese IP and the Phillips–Perron $Z(t\alpha^{\sim})$ statistic of the British TB reject the unit root but other statistics, including the unreported Phillips–Perron $Z(\Phi_1)$, $Z(\Phi_2)$, $Z(\Phi_3)$ statistics, of the two variables fail to reject the hypothesis. The hypothesis is overwhelmingly rejected in both tests for all the first–differenced series. These results indicate that all the series of the five countries appear to be I(1) and first–differencing will likely make them stationary.

3) Schwert argues that the critical values of the Phillips–Perron are far below the Dickey–Fuller critical values and that the best in the sense that it is least affected by the different processes used to generate the data is the ADF test.

Table 3. Unit root test for level and first−differenced data

		s	Δs	m	Δm	ip	Δip
	$t\alpha'$	-0.94(6)	-5.55*(6)	3.47(12)	-2.65*(12)	3.56(1)	-11.79*(1)
Canada	$t\alpha^{*}$	-1.60(6)	-5.63*(6)	-2.44(12)	-4.19*(12)	-1.14(1)	-12.70*(1)
	$t\alpha^{\sim}$	-0.98(6)	-5.83*(6)	-1.99(12)	-4.57*(12)	-2.18(1)	-12.72*(1)
	$Z(t\alpha')$	-1.02	-14.9*	2.74	-19.1*	2.65	-19.4*
	$Z(t\alpha^{*})$	-1.51	-15.0*	-1.69	-20.0*	-0.69	-20.7*
	$Z(t\alpha^{\sim})$	-0.86	-6.83*	-3.18	-5.75*	-2.72	-5.61*
	$t\alpha'$	0.84(1)	-9.87*(1)	-0.46(6)	-4.83*(6)	2.04(2)	-10.3*(2)
Germany	$t\alpha^{*}$	-1.36(1)	-9.86*(1)	-0.46(6)	-4.83*(6)	0.52(2)	-10.6*(2)
	$t\alpha^{\sim}$	-1.67(1)	-9.85*(1)	-2.21(6)	-4.81*(6)	-2.13(2)	-10.7*(2)
	$Z(t\alpha')$	0.91	-12.8*	3.05	-13.6*	1.58	-26.8*
	$Z(t\alpha^{*})$	-1.47	-12.9*	-0.32	-14.3*	0.01	-27.2*
	$Z(t\alpha^{\sim})$	-1.73	-9.33*	-3.31	-9.18*	-2.71	-4.03*
	$t\alpha'$	1.02(12)	-3.56*(12)	2.98(12)	-1.97*(12)	1.82(12)	-4.09*(12)
Japan	$t\alpha^{*}$	-0.94(12)	-3.77*(12)	-1.78(12)	-3.24*(12)	-0.66(12)	-4.77*(12)
	$t\alpha^{\sim}$	-2.85(12)	-3.78*(12)	-2.44(12)	-3.47*(12)	-4.71*(12)	-4.68*(12)
	$Z(t\alpha')$	1.36	-12.4*	3.18	-21.8*	2.75	-22.1*
	$Z(t\alpha^{*})$	-0.35	-12.6*	-1.34	-22.7*	-0.02	-23.4*
	$Z(t\alpha^{\sim})$	-2.12	-9.58*	-3.19	-3.80*	-2.56	-3.54*
	$t\alpha'$	-0.40(1)	-10.3*(1)	1.68(12)	-1.97*(12)	-0.83(4)	-4.08*(12)
France	$t\alpha^{*}$	-1.31(1)	-10.4*(1)	-2.68(12)	-3.24*(12)	-0.84(4)	-4.77*(12)
	$t\alpha^{\sim}$	0.51(1)	-10.4*(1)	0.17(12)	-3.47*(12)	-2.83(4)	-4.67*(12)
	$Z(t\alpha')$	-0.32	-14.0*	4.21	-20.6*	1.39	-29.1*
	$Z(t\alpha^{*})$	-1.28	-14.0*	-2.39	-21.2*	-1.25	-29.3*
	$Z(t\alpha^{\sim})$	-0.87	-7.18*	-0.46	-10.8*	-2.94	-3.99*
	$t\alpha'$	-0.64(1)	-9.71*(1)	-0.92(6)	-7.84*(5)	0.61(4)	-6.78*(4)
U.K.	$t\alpha^{*}$	-2.12(1)	-9.71*(1)	-0.92(6)	-4.67*(5)	-1.12(4)	-6.79*(4)
	$t\alpha^{\sim}$	-1.63(1)	-9.81*(1)	-0.97(6)	-4.74*(5)	-2.64(4)	-6.79*(4)
	$Z(t\alpha')$	-0.58	-12.5*	9.33	-6.97*	0.64	-18.8*
	$Z(t\alpha^{*})$	-1.99	-12.5*	-1.74	-12.9*	-1.31	-18.8*
	$t\alpha'$	-1.55	-11.5*	-0.78	-9.43*	-2.61	-5.54*
	$t\alpha'$			2.16(6)	-7.69*(4)	1.63(1)	-6.42*(1)
U.S.	$t\alpha^{*}$			-0.12(6)	-5.51*(4)	-0.66(1)	-6.58*(1)
	$t\alpha^{\sim}$			-1.98(6)	-5.49*(4)	-2.62(1)	-6.57*(1)
	$Z(t\alpha')$			6.02	-15.6*	2.20	-6.65*
	$Z(t\alpha^{*})$			-0.01	-19.4*	-0.45	-6.84*
	$Z(t\alpha^{\sim})$			-2.45	-4.16*	-2.07	-7.29*

Table 3. (Continued)

		I_S	ΔI_S	I_L	ΔI_L	ΔTB	TB
	$t\alpha'$	-0.69(2)	-7.86*(2)	-0.18(2)	-8.34*(2)	-1.23(6)	-6.59*(6)
Canada	$t\alpha^*$	-2.59(2)	-7.84*(2)	-1.98(2)	-8.33*(2)	-1.23(6)	-6.60*(6)
	$t\alpha^{\sim}$	-2.41(2)	-7.92*(2)	-1.80(2)	-8.39*(2)	-0.99(6)	-6.62*(6)
	$Z(t\alpha')$	-0.52	-8.48*	-0.16	-12.4*	-1.10	-23.1*
	$Z(t\alpha^*)$	-2.18	-8.48*	-2.07	-12.4*	-2.43	-23.1*
	$Z(t\alpha^{\sim})$	-1.93	-9.49*	-1.88	-10.5*	-2.51	-4.75*
	$t\alpha'$	-1.25(1)	-8.47*(1)	-0.69(1)	-8.93*(1)	-1.81(7)	-4.41*(6)
Germany	$t\alpha^*$	-2.39(1)	-8.47*(1)	-2.26(1)	-8.92*(1)	-1.81(7)	-4.41*(6)
	$t\alpha^{\sim}$	-2.24(1)	-8.70*(1)	-2.17(1)	-8.95*(1)	-2.26(7)	-4.54*(6)
	$Z(t\alpha')$	-0.73	-8.92*	-0.48	-8.05*	-1.39	-31.1*
	$Z(t\alpha^*)$	-1.67	-8.93*	-1.73	-8.06*	-2.35	-31.1*
	$Z(t\alpha^{\sim})$	-1.54	-13.4*	-1.67	-8.66*	-2.53	-4.21*
	$t\alpha'$	-1.48(13)	-3.99*(12)	-0.71(4)	-7.85*(3)	-1.02(12)	-3.20*(12)
Japan	$t\alpha^*$	-2.07(13)	-4.02*(12)	-1.34(4)	-7.84*(3)	-1.03(12)	-3.13*(12)
	$t\alpha^{\sim}$	-3.22(13)	-4.04*(12)	-2.30(4)	-7.85*(3)	-2.62(12)	-3.45*(12)
	$Z(t\alpha')$	-0.69	-8.25*	-0.65	-11.0*	-1.10	-44.2*
	$Z(t\alpha^*)$	-1.64	-8.25*	-1.42	-11.0*	-2.21	-50.1*
	$Z(t\alpha^{\sim})$	-2.21	-12.2*	-2.42	-37.5*	-2.36	-6.63*
	$t\alpha'$	-0.48(1)	-3.99*(12)	-0.22(1)	-9.99*(1)	-0.06(11)	-3.60*(12)
France	$t\alpha^*$	-2.64(1)	-4.02*(12)	-1.27(1)	-9.97*(1)	-2.27(11)	-3.81*(12)
	$t\alpha^{\sim}$	-2.70(1)	-4.04*(12)	-1.61(1)	-9.97*(1)	-2.21(11)	-3.85*(12)
	$Z(t\alpha')$	-0.35	-8.71*	-0.21	-11.1*	-1.21	-35.8*
	$Z(t\alpha^*)$	-2.31	-8.71*	-1.25	-11.1*	-2.30	-35.8*
	$Z(t\alpha^{\sim})$	-2.38	-12.3*	-1.59	-11.5*	-3.24	-4.52*
	$t\alpha'$	-0.42(1)	-9.84*(1)	-0.60(6)	-6.78*(4)	0.34(8)	-7.62*(6)
U.K.	$t\alpha^*$	-2.21(1)	-9.82*(1)	-1.77(6)	-6.77*(4)	-1.21(8)	-6.29*(6)
	$t\alpha^{\sim}$	-3.13(1)	-9.84*(1)	-2.98(6)	-6.79*(4)	-2.01(8)	-6.28*(6)
	$Z(t\alpha')$	-0.45	-12.8*	-0.47	-14.8*	-1.14	-36.5*
	$Z(t\alpha^*)$	-2.09	-12.8*	-1.81	-14.8*	-2.53	-36.5*
	$Z(t\alpha^{\sim})$	-3.02	-15.1*	-3.28	-9.98*	-3.44*	-4.81*
	$t\alpha'$	-0.72(6)	-8.48*(5)	-0.16(4)	-8.76*(2)	-0.15(8)	-7.34*(6)
U.S.	$t\alpha^*$	-1.26(6)	-8.48*(5)	-1.54(4)	-8.75*(2)	-1.25(8)	-7.54*(6)
	$t\alpha^{\sim}$	-1.33(6)	-8.54*(5)	-1.26(4)	-8.84*(2)	-0.99(8)	-7.57*(6)
	$Z(t\alpha')$	-0.85	-9.35*	-0.12	-11.2*	-0.21	-28.5*
	$Z(t\alpha^*)$	-2.13	-9.36*	-1.80	-11.2*	-1.57	-28.6*
	$Z(t\alpha^{\sim})$	-2.20	-10.7*	-1.51	-11.6*	-2.84	-3.45*

Notes: 1. S denotes spot exchange rate, m money supply, ip industrial production, i_S short-term interest rate, i_L long-term interest rate and TB trade balance.

2. $t\alpha'$, $t\alpha^*$, $t\alpha^{\sim}$ = Augumented Dickey−Fuller test statistic with the number of lags in parentheses.

3. $Z(t\alpha^{\wedge})$, $Z(t\alpha^{*})$, $Z(t\alpha^{\sim})$ = Phillips−Perron test statistic for a unit root. Number of truncation lags in its construction is 2. When 4 lags and 8 lags were used similar, qualitatively, results were obtained.
4. The critical values of the $t\alpha^{\wedge}(=Z(t\alpha^{\wedge}))$, $t\alpha^{*}(=Z(t\alpha^{*}))$, and $t\alpha^{\sim}$ $(=Z(t\alpha^{\sim}))$ at the 5 percent significance level with sample of size 250 is −1.95 (−2.58), −2.88(−3.46) and −3.43(−3.99) respectively. The critical values for the unit root tests are tabulated in Fuller(1976, p.373)
5. *Significant at the 5% level.

4. EMPIRICAL EVIDENCE

We now proceed test for cointegration using the maximum likelihood methodology of Johansen (1988) and Johansen and Juselius (1990). The results are reported in Table 4.

Table 4. Test of Cointegration

Frenkel−Bilson model:

trace	Can.	Ger.	Jap.	Fra.	U.K.	λ_{max}		Can.	Ger.	Jap.	Fra.	U.K.
r≤5	5.024	4.661	9.187	15.48	4.828	r=5	r=6	4.346	4.656	7.569	12.91	4.043
r≤4	18.04	19.35	17.76	30.35	14.78	r=4	r=5	13.92	14.69	8.575	14.87	9.954
r≤3	40.11	37.50	39.52	49.55	30.78	r=3	r=4	21.16	18.15	21.76	19.19	16.00
r≤2	70.78	63.51	70.28	72.17	56.59	r=2	r=3	30.67	26.01	30.75	28.62	25.81
r≤1	112.8	117.0	116.0	114.1	85.72	r=1	r=2	42.09	53.54	45.74	35.99	29.13
r=0	172.6	171.8	166.5	172.9	133.2	r=0	r=1	59.78	54.82	50.55	58.73	47.52

Dornbusch−Frankel model:

trace	Can.	Ger.	Jap.	Fra.	U.K.	λ_{max}		Can.	Ger.	Jap.	Fra.	U.K.
r≤8	0.306	0.015	2.221	2.282	0.945	r=8	r=9	0.306	0.015	2.221	2.282	0.945
r≤7	10.14	8.031	9.711	12.45	5.962	r=7	r=8	9.842	8.015	7.489	10.17	5.017
r≤6	25.33	25.44	18.81	28.99	17.80	r=6	r=7	15.18	17.41	9.100	16.53	11.84
r≤5	43.90	49.41	43.71	49.74	35.46	r=5	r=6	18.56	23.96	24.90	20.74	17.65
r≤4	65.88	75.70	71.29	84.72	57.71	r=4	r=5	21.92	26.28	27.57	34.98	22.25
r≤3	96.92	104.9	103.9	126.9	85.03	r=3	r=4	31.10	29.26	32.68	42.17	27.32
r≤2	143.9	142.4	156.9	175.7	121.6	r=2	r=3	46.98	37.44	52.96	48.85	36.57
r≤1	195.1	201.2	216.5	229.9	171.7	r=1	r=2	51.23	58.88	59.64	54.19	50.17
r=0	269.2	291.5	288.9	297.9	239.2	r=0	r=1	74.09	90.23	72.32	68.01	67.49

Hooper−Morton model:

trace	Can.	Ger.	Jap.	Fra.	U.K.	λ_{max}		Can.	Ger.	Jap.	Fra.	U.K.
r≤10	0.837	1.016	1.987	1.038	0.023	r=10	r=11	0.837	1.016	1.987	1.038	0.023
r≤9	7.895	8.162	9.364	10.16	5.360	r=9	r=10	7.811	7.145	7.376	9.131	5.336
r≤8	20.40	23.74	19.19	26.49	18.75	r=8	r=9	12.51	15.58	9.829	16.32	13.39
r≤7	38.79	42.59	40.77	48.15	36.28	r=7	r=8	18.38	18.85	21.57	21.66	17.52
r≤6	62.09	71.68	65.94	71.08	55.56	r=6	r=7	23.30	29.09	25.17	22.93	19.27
r≤5	94.98	105.7	95.85	100.4	88.24	r=5	r=6	32.88	34.02	29.91	29.31	32.68
r≤4	133.7	147.1	131.1	138.6	126.9	r=4	r=5	38.73	41.38	35.29	38.27	38.69
r≤3	177.0	203.3	181.8	192.8	169.1	r=3	r=4	43.30	56.26	50.20	54.20	42.23
r≤2	227.8	263.5	242.8	256.4	223.1	r=2	r=3	50.73	60.18	61.50	63.57	53.96
r≤1	295.9	347.2	330.4	324.5	291.5	r=1	r=2	68.16	88.73	87.58	68.09	68.39
r=0	410.4	471.3	432.9	413.9	371.4	r=0	r=1	114.4	124.1	102.5	89.37	79.93

Notes: 1. r is the number of distinct cointegrating vectors.
2. Trace is the trace statistic and λ_{max} is the maximal eigenvalue statistic.
3. The critical values are obtained from the Osterwald−Lenum (1992).

Before the likelihood−ratio test statistic can be computed, the order of the lag length in equation (1) must be determined. This is done by testing down from a general VAR to the minimum number of significant lags using standard likelihood−ratio tests with the degrees of freedom correction recommended by Sims(1980). A test is then made for misspecification in the chosen VARs by examining the whiteness of the residuals using Ljung−Box portmanteau statistics. If the residuals in any equation proved to be non−white we sequentially choose a higher lag structure until they were whitened. For all three models we found that a eighth−order lag satisfied this criterion.

The results are reported in Table 4. In all of the three models, more than one cointegrating vectors are detected. In the Frenkel−Bilson model, we consider the number of cointegrating vectors, beginning with the hypothesis $r \leq 2$ against the general unrestricted alternative $r = p$. The trace statistics for Canada, Germany, Japan, France and U.K. are 70.78, 63.51, 70.28, 72.17, and 56.59 respectively, all of which fall short of the ctitical value 76.07 at 5% significance level. However, the trace statistics for $r \leq 1$ hypothesis are 112.8, 117.0, 116.0 and 114.1 respectively, which exceed the critical value 102.14 at 5 % significance level, except U.K. This indicates that there are at most two cointegrating vectors for Canada, Germany, Japan and France, but that one vector for U.K. in the Frenkel−Bilson model. The maximal eigenvalue statistics support this finding. The λ_{MAX} statistics for $r = 0$ hypothesis within the hypothesis of $r = 1$ are 59.78, 54.82, 50.55, 58.73 and 47.52 respectively for five countries and these all exceed the critical value 46.45 at 5 % level. This indicates that there exists at least one cointegrating vector. This critical values for both trace statistic and maximal eigenvalue statistic are obtained from Osterwald−Lenum (1992), who extended the tables presented in Johansen (1988) and Johansen and Juselius (1990) up to 11 variables.

Similarly, we find that there exist at most 3,4,4,5 and 2 cointegrating vectors for Canada, Germany, Japan, France and U.K. respectively in the Dornbusch−Frankel model. In the Hooper−Morton model, there exist at most 5, 6, 4, 5 and 4 cointegrating vectors respectively for the five countries. This suggests that the long−run relationship is a useful concept for exchange rate modeling for the five countries.

Our results indicate a cointegrating relationship between the exchange rates and the monetary variables for Canada, Germany, Japan, France and the U.K. A cointegrating relationship among the series of I(1) can be represented by an error−correction model. The estimated error−correction equations for the three models are reported in Appendix.

Next, the long−run forecasting is made for the three models based on these estimated error−correction equations. Before proceeding to the long−run forecast, we report the diagnostic tests for the error−correction models used for long−run forecasting. The Ljung−Box Q statistic and the Lagrange Multiplier test indicate non−rejeciton of the null hypothesis of no serial correlation. The ARCH statistics which test for conditional heteroskedasticity fail in rejecting the null hypothesis of homoskedasticity. The HETERO statistic which is White's test for heteroskedasticity also indicates non−rejection. The CHOW statistics for structural stability show non−rejection of the null hypothesis of stability. These evidences support the use of the error−correction model for the purpose of long−run forecasting.

Table 5. Diagnostic test for the error−correction model.

model		R^2	Q	LM	ARCH	CHOW	HETERO
	F−B	0.28	49.72	9.161	6.826	0.684	19.71
			(0.290)	(0.329)	(0.555)	(0.723)	(0.874)
Canada	D−F	0.38	48.03	9.552	4.987	1.398	7.554
			(0.282)	(0.297)	(0.758)	(0.235)	(0.994)
	H−M	0.44	47.73	10.49	6.262	0.794	12.02
			(0.298)	(0.232)	(0.618)	(0.575)	(0.915)
	F−B	0.20	48.67	6.084	3.975	0.782	12.81
			(0.328)	(0.638)	(0.859)	(0.585)	(0.885)
France	D−F	0.33	53.24	1.866	3.638	0.882	12.64
			(0.187)	(0.985)	(0.888)	(0.508)	(0.892)
	H−M	0.39	52.18	4.761	4.253	1.439	27.39
			(0.216)	(0.783)	(0.833)	(0.172)	(0.816)

	F−B	0.23	45.33	4.356	3.843	1.061	17.70
			(0.459)	(0.823)	(0.999)	(0.389)	(0.912)
Germany	D−F	0.34	39.11	11.82	2.211	1.827	10.88
			(0.718)	(0.159)	(0.920)	(0.108)	(0.695)
	H−M	0.42	43.24	8.839	2.057	1.601	13.84
			(0.546)	(0.355)	(0.979)	(0.147)	(0.792)
	F−B	0.21	47.06	3.984	14.98	0.966	11.53
			(0.387)	(0.858)	(0.132)	(0.426)	(0.241)
Japan	D−F	0.34	44.59	3.757	14.33	1.162	28.15
			(0.491)	(0.878)	(0.127)	(0.325)	(0.402)
	H−M	0.40	40.12	3.953	15.44	1.063	25.05
			(0.678)	(0.861)	(0.116)	(0.390)	(0.571)
	F−B	0.28	47.24	3.768	5.810	0.868	13.21
			(0.382)	(0.877)	(0.668)	(0.503)	(0.794)
U.K.	D−F	0.34	37.03	3.938	8.696	0.404	12.04
			(0.794)	(0.862)	(0.368)	(0.845)	(0.915)
	H−M	0.45	45.60	5.017	8.509	1.128	22.18
			(0.446)	(0.755)	(0.385)	(0.346)	(0.727)

Notes. 1. LM is the Chi−square statistic in a Lagrange Multiplier test for eighth order serial correlation; ARCH is the Chi−square statistic for eighth order conditional heteroscedasticity; CHOW is the F test for structural stability; and HETERO is a general test of heteroscedasticity.

2. R_2 is the coefficient of determination; SEE is the standard error of the regression; The Q statistic is the value of the Ljung−Box test for autocorrelated residuals; and marginal significance levels are reported beneath each statistic.

Forecasting performance is compared between the structural monetary models and the random walk model. Most of the previous studies on exchange rate forecasting have been concerned with the short−run forecasting of the structural models. We present the results of short−run and long−run out−of−sample forecasts of the three monetary models and the random walk model. The out−of−sample forecasts are made by dynamic rolling regression. The structural models in this paper are initially estimated using data through the first forecasting period, 1990:6. Forecasts are then generated at the one−, two−, three−, six−, nine−, and twelve−month horizons. Next, the data for July 1990 are added to the original data set and each model is re−estimated. New forecasts are made for the six time horizons. This "rolling regression" process is continued through to the last forecast period, 1993:6. Forecasting performance is measured by three statistics: root mean squared error (RMSE), mean absolute error

(MAE) and mean error (ME). They are defined as follows.

$$\mathrm{RMSE} = \left(\sum_{s=0}^{N_{k-1}} [F(t+s+k) - A(t+s+k)]^2 / N_k \right)^{1/2}$$

$$\mathrm{MAE} = \sum_{s=0}^{N_{k-1}} | F(t+s+k) - A(t+s+k) | / N_k$$

$$\mathrm{ME} = \sum_{s=0}^{N_{k-1}} [F(t+s+k) - A(t+s+k)] / N_k$$

where k=1, 2, 3, 6, 9, 12 (the forecast step); F(p) and A(p) denote the forecasted and actual values of the exchange rate for period 'p'; Nk is the number of forecasts throughout the entire forecasting period; and t starting period for forecasting. Following Meese and Rogoff (1983), RMSE is our principal criterion for comparing models' forecasts. An advantage of the RMSE measure is that it penalizes more relatively large forecast errors. The RMSE criterion is inappropriate if, for example, the exchange rate follows a non−normal stable Paretian process with infinite variance, as suggested by Westerfield (1977). The stable Paretian family, of which the normal distribution is a special case, is characterized by 'fat tails'. Westerfield shows that the normal model is a special probability model where the distributions are symmetric and completely described by their means and variances. The major implication of a nonnormal stable model is that means and variances do not adequately describe the probabilistic properties of foreign exchange rates. In that case MAE, which is robust to fat−tailed distribution, will be a very useful criterion. The ME statistic is also reported as it is useful for check of a systematic bias of the models.

Table 6. Forecast errors of the stuctural models: RMSE

		short−run					long−run				
model	step	Can	Fra	Ger	Jap	U.K.	Can	Fra	Ger	Jap	U.K.
F−B	1	10.79	15.99	10.79	7.896	21.87	1.075	3.921	3.685	3.551	4.208
model	2	11.26	16.77	11.61	8.291	22.94	1.072	3.895	3.849	3.497	4.351
	3	11.62	17.56	12.47	8.632	23.98	1.100	3.933	3.889	3.539	4.395
	6	12.54	19.45	15.22	8.367	26.91	1.149	4.080	4.005	3.411	4.493
	9	13.25	20.81	17.27	7.016	30.15	1.156	4.266	4.255	3.546	4.468
	12	14.03	23.18	17.41	6.787	31.70	1.181	3.884	3.739	3.455	4.771

D−F model	1	7.748	14.49	8.089	8.457	15.01	1.214	3.914	3.944	3.539	4.205
	2	8.156	15.35	8.812	9.092	15.91	1.214	3.867	3.938	3.546	4.335
	3	8.443	16.19	9.447	9.024	15.75	1.234	3.907	3.982	3.593	4.379
	6	9.264	18.76	11.03	6.863	19.19	1.284	4.063	4.085	3.430	4.480
	9	10.00	21.16	11.56	7.054	21.76	1.305	4.235	4.217	3.563	4.657
	12	10.58	23.72	10.94	7.387	23.13	1.366	3.821	3.754	3.416	4.754
H−M model	1	5.056	14.05	8.145	8.775	15.49	1.226	3.922	3.890	3.488	4.071
	2	5.216	14.89	9.036	9.331	16.51	1.225	3.903	3.896	3.480	4.133
	3	5.249	15.67	9.750	9.304	17.51	1.247	3.946	3.938	3.518	4.170
	6	5.209	18.20	11.43	6.544	20.40	1.298	4.088	4.046	3.387	4.269
	9	5.626	20.53	12.06	6.745	23.38	1.332	4.253	4.290	3.522	4.407
	12	6.315	23.07	11.21	7.630	25.02	1.399	3.845	3.721	3.443	4.479

Table 7. Random walk forecast errors

		With No Drift			With Drift		
	step	RMSE	MAE	ME	RMSE	MAE	ME
Canada	1	1.238	0.912	−0.179	1.241	0.910	−0.193
	2	1.173	1.335	−0.383	1.738	1.338	−0.405
	3	2.292	1.749	−0.598	2.295	1.756	−0.625
	6	3.155	2.504	−1.221	3.166	2.549	−1.235
	9	4.083	3.637	−2.142	4.098	3.672	−2.084
	12	5.201	4.378	−2.846	5.108	4.330	−2.720
France	1	3.943	2.985	0.045	3.924	2.980	0.075
	2	5.905	4.552	0.105	5.846	4.522	0.064
	3	7.202	5.816	0.157	7.094	5.752	0.232
	6	9.909	8.214	−0.251	9.595	7.961	−0.084
	9	10.17	9.201	−0.609	9.668	8.721	−0.382
	12	10.33	8.553	1.771	9.771	8.051	1.886
Germany	1	3.992	2.909	−0.277	3.986	2.929	0.104
	2	6.051	4.601	−0.525	6.040	4.713	0.236
	3	7.523	5.960	−0.812	7.495	5.997	0.327
	6	10.29	8.178	−2.128	10.05	8.364	0.140
	9	10.48	8.913	−3.545	9.817	8.813	−0.172
	12	9.960	8.352	−2.167	9.783	8.084	2.290
Japan	1	3.540	2.679	0.612	3.591	2.699	0.820
	2	4.225	3.514	0.861	4.349	3.268	1.279
	3	5.165	4.299	1.296	5.390	4.460	1.932
	6	5.887	4.921	1.464	6.395	5.260	2.761
	9	4.572	3.627	1.421	5.566	4.593	3.391
	12	5.380	4.212	2.330	7.052	5.675	5.014

U.K.	1	4.311	3.137	0.130	4.268	3.099	0.106
	2	6.606	4.920	0.861	6.456	4.807	0.591
	3	7.572	6.087	1.296	7.322	5.888	1.321
	6	10.02	8.649	1.464	9.393	8.056	2.816
	9	12.35	11.31	1.421	11.21	10.18	3.767
	12	12.29	10.47	2.330	10.77	9.004	3.856

Table 8. Short-run forecast errors: MAE and ME

		Canada		France		Germany		Japan		U.K.	
model	step	MAE	ME	MAE	ME	MAE	ME	MAE	ME	MAE	ME
F-B	1	9.645	9.361	13.11	12.83	9.028	-4.325	6.339	3.273	20.45	19.91
model	2	9.998	9.719	13.73	13.48	9.650	-4.509	6.707	3.501	21.61	21.06
	3	10.29	10.00	14.44	14.21	10.33	-4.733	6.986	3.669	22.72	22.24
	6	10.99	10.76	15.93	15.89	12.55	-5.496	6.883	3.146	25.65	25.42
	9	11.52	11.34	17.10	17.10	14.41	-5.746	6.317	2.101	29.12	29.08
	12	12.41	12.26	19.88	19.88	14.66	-5.176	6.191	1.530	30.69	30.69
D-F	1	6.835	5.964	12.22	11.53	6.813	-3.937	6.468	5.681	13.54	12.42
model	2	7.156	6.225	12.99	12.31	7.488	-4.367	6.809	6.018	14.51	13.51
	3	7.385	6.398	13.75	13.12	8.125	-4.986	6.761	5.965	15.45	14.62
	6	7.995	6.966	16.01	15.60	9.616	-6.651	5.954	5.169	18.06	17.42
	9	8.403	7.514	18.10	17.83	10.12	-7.196	6.151	5.470	21.04	20.74
	12	8.770	8.156	20.71	20.71	9.840	-7.058	6.397	5.768	22.34	22.22
H-M	1	4.438	2.277	11.79	10.77	6.798	-3.072	6.180	2.144	13.80	12.94
model	2	4.565	2.283	12.54	11.52	7.636	-3.598	6.416	2.043	14.97	14.04
	3	4.575	2.234	13.29	12.30	8.257	-4.306	6.285	1.676	15.93	15.19
	6	4.532	2.086	15.49	14.79	9.671	-6.278	5.057	-0.271	19.06	18.52
	9	4.867	2.290	17.36	16.94	10.41	-6.905	5.189	-0.681	22.54	22.17
	12	5.399	2.820	19.94	19.84	9.902	-6.615	5.698	-0.956	42.16	23.91

Table 9. Long-run forecast errors: MAE and ME

		Canada		France		Germany		Japan		U.K.	
model	step	MAE	ME	MAE	ME	MAE	ME	MAE	ME	MAE	ME
F-B	1	0.801	-0.090	2.974	0.077	2.750	-0.028	2.670	0.323	3.085	-0.273
model	2	0.791	-0.122	2.918	-0.066	2.694	-0.108	2.650	0.234	3.183	-0.242
	3	0.814	-0.121	2.914	-0.129	2.716	-0.171	2.680	0.196	3.209	-0.312
	6	0.868	-0.122	3.103	-0.258	2.823	-0.310	2.487	0.007	3.272	-0.380
	9	0.869	-0.225	3.281	-0.237	3.148	-0.293	2.591	0.018	3.396	-0.572
	12	0.868	-0.316	3.193	0.215	2.960	0.175	2.523	0.253	3.421	-0.881

D-F model	1	0.889	-0.156	2.901	-0.055	2.837	-0.215	2.589	0.437	3.084	-0.217
	2	0.879	-0.196	2.811	-0.199	2.801	-0.355	2.582	0.369	3.186	-0.178
	3	0.900	-0.199	2.833	-0.264	2.829	-0.427	2.624	0.339	3.209	-0.241
	6	0.957	-0.210	2.999	-0.395	2.940	-0.546	2.408	0.155	3.267	-0.290
	9	0.974	-0.284	3.144	-0.365	3.148	-0.525	2.486	0.211	3.398	-0.457
	12	1.015	-0.351	3.050	0.063	2.979	-0.053	2.386	0.481	3.417	-0.753
H-M model	1	0.909	-0.188	2.958	0.005	2.690	-0.306	2.644	0.486	3.054	-0.295
	2	0.898	-0.230	2.908	-0.126	2.650	-0.441	2.621	0.407	3.092	-0.235
	3	0.920	-0.233	2.931	-0.187	2.670	-0.509	2.645	0.364	3.092	-0.322
	6	0.976	-0.241	3.091	-0.323	2.763	-0.650	2.457	0.186	3.162	-0.393
	9	1.004	-0.306	3.239	-0.296	3.094	-0.601	2.579	0.208	3.226	-0.561
	12	1.050	-0.373	3.106	0.072	2.912	-0.131	2.498	0.461	3.225	-0.859

Table6, Table 8 and Table 9 present the short-run and long-run forecasting errors of the three monetary models for the period of 90:7 through 93:6. Table 7 shows the forecast errors of the random walk model of exchange rates. The finding of Meese and Rogoff is largely confirmed: The RMSE and MAE statistics indicate that the random walk model outperforms the short-run monetary models in all forecasting steps for all countries. The only exception is the 12-step forecast of the Frenkel-Bilson model for Japan, beating the random walk with drift. However, the ME statistic for random walk with drift for Japan indicates a systematic forecasting bias. The ME statistics strongly indicate systematic bias in the short-run structural models: They are F-B and D-F models for Canada, D-F model for Japan and all three models for France and U.K. A less significant bias is shown in the D-F and H-M models for Germany. The ME statistic indicates no systematic bias for the random walk model for all countries except Japan. There is no significant difference in the forecasting performance between the random walk with drift and that without drift.

The results of the long-run out-of-sample forecasts are markedly different from those of the short-run model. The tables show that for the three criteria, all three structural models overwhelmingly outperform the random walk model at all steps for all five countries. And we find that the long-run forecasts of the structural models are very stable compared with those of the random walk model.

The ME statistic indicates no systematic bias in the long-run models for all forecast steps, for all structural models and for all countries. This is an empirical confirmation of the survey findings of Frankel and Froot (1990) and Taylor and Allen (1992) that while the dealers in the foreign exchange markets rely more on charts and recent trends at short horizons, they tend to forecast rates of return more based on the economic fundamentals at long horizons.

5. CONCLUDING REMARKS

This paper reexamined the monetary approach to the exchange rate determination from a new perspective. In particular, a maximum likelihood methodology for estimating the cointegrating vectors, proposed by Johansen (1988) and Johansen and Juselius (1990), has been implemented in order to test for the validity of the long−run monetary models. We used monthly data for for Canada, France, Germany, Japan, U.K. and U.S. for the period of 1973:4 through 1993:6.

We found a long−run relationship for the the Frenkel−Bilson model, the Dornbusch−Frankel model, and the Hooper−Morton model. Our finding of the cointegrating vector indicates that the monetary model can be interpreted as having long−run validity. We proceeded by comparing the forcasting performance of the three structural models of exchange rates with that of the random walk model (with and without drift) for the short−run and the long−run using RMSE, MAE and ME statistics. The long−run forecasted values of the exchange rates were computed using the dynamic error correction equations. The result was that while the structural monetary models provide poorer performance than the random walk for short−run forecasting, they outperform the random walk model in the long−run forecasting. Our finding leads to the conclusion that the monetary models shoud be thought more of as the long−run models.

APPENDIX. Error−correction model

Frenkel−Bilson model:

Canada $R_2 = 0.28$, $Q(45) = 49.72(0.290)$

$$\Delta S_t = \underset{(0.080)}{-0.138}\ \Delta S_{t-7} \underset{(0.098)}{-0.214}\ \Delta IP_{t-2} \underset{(0.001)}{+0.004}\ \Delta SHORT_{t-1} \underset{(0.002)}{+0.004}\ \Delta SHORT_{t-3}$$

$$\underset{(0.002)}{-0.005}\ \Delta SHORT_{t-4} \underset{(0.002)}{-0.006}\ \Delta SHORT_{t-7} \underset{(0.002)}{-0.005}\ \Delta SHORT^{*}_{t-3}$$

$$\underset{(0.017)}{-0.040}\ \Delta SHORT^{*}_{t-5}$$

France $R^2 = 0.20$, $Q(45) = 48.65(0.328)$

$$\Delta S_t = \underset{(0.080)}{-0.161}\ \Delta S_{t-6} \underset{(0.023)}{-0.049}\ \Delta SHORT^{*}_{t-8}$$

Germany $R^2 = 0.23$, $Q(45) = 45.31(0.459)$

$$\Delta S_t = \underset{(0.424)}{-1.244} \underset{(0.082)}{-0.148}\ \Delta S_{t-3} \underset{(0.080)}{-0.229}\ \Delta S_{t-6} \underset{(0.180)}{+0.491}\ \Delta IP_{t-6} \underset{(0.190)}{+0.389}\ \Delta IP_{t-7}$$

$$- 0.105\ M_{t-5} + 0.443\ IP_{t-5} - 0.005\ SHORT_{t-5}$$
(0.045) (0.174) (0.002)

Japan $R^2 = 0.21$, $Q(45) = 47.06(0.387)$

$$\Delta S_t = -0.702\ \Delta IP^*_{t-1} + -0.069\ S_{t-8} - 0.003\ SHORT^*_{t-5}$$
(0.381) (0.026) (0.001)

U.K. $R^2 = 0.28$, $Q(45) = 47.21(0.382)$

$$\Delta S_t = -7.333 + 0.430\ \Delta M_{t-1} + 0.424\ \Delta M^*_{t-7}$$
(0.413) (0.195) (0.219)

$$-0.053\ S_{t-8} + 0.302\ \Delta IP_{t-5} + 0.003\ SHORT_{t-5}$$
(0.023) (0.125) (0.001)

Dornbusch− Frankel model

Canada $R^2 = 0.38$, $Q(45) = 48.02(0.282)$

$$\Delta S_t = 0.487 - 0.184\ \Delta DS_{t-7} - 0.081\ S_{t-5} + 0.045\ M^*_{t-5}$$
(0.246) (0.086) (0.022) (0.022)

France $R^2 = 0.33$, $Q(45) = 53.21(0.187)$

$$\Delta S_t = -0.170\ \Delta S_{t-1} - 0.183\ \Delta S_{t-6} - 0.025\ LONG^{**}_{t-1}$$
(0.082) (0.083) (0.011)

$$-0.079\ S_{t-8} - 0.018\ LONG^*_{t-5}$$
(0.031) (0.008)

Germany $R^2 = 0.34$, $Q(45) = 39.11(0.718)$

$$\Delta S_t = -0.198\ \Delta S_{t-4} - 0.153\ \Delta S_{t-7} - 0.075\ S_{t-5} + 0.408\ IP_{t-5}$$
(0.078) (0.078) (0.032) (0.141)

Japan $R^2 = 0.34$, $Q(45) = 44.52(0.491)$

$$\Delta S_t = -0.192\ \Delta S_{t-7} - 0.029\ \Delta LONG_{t-7} - 0.036\ \Delta LONG^*_{t-1} + 0.028\ \Delta LONG^*_{t-7}$$
(0.090) (0.012) (0.011) (0.012)

$$-0.138\ S_{t-5} - 0.012\ LONG^*_{t-5}$$
(0.044) (0.005)

U.K. $R^2 = 0.34$, $Q(45) = 37.03(0.794)$

$$\Delta S_t = 0.418\ \Delta M_{t-1} + 0.383\ \Delta IP_{t-3} + 0.008\ \Delta SHORT_{t-1}$$
(0.202) (0.198) (0.004)

$$-0.065\ S_{t-5} + 0.368\ IP_{t-5}$$
(0.003) (0.160)

Hooper−Morton model:

Canada $R^2 = 0.44$, $Q(45) = 47.75(0.298)$

$$\Delta St = 0.529 - 0.200\ \Delta S_{t-2} - 0.189\ \Delta S_{t-7} + 0.230\ \Delta IP^*_{t-1}$$
(0.253) (0.095) (0.092) (0.148)

$$+0.005\ \Delta SHORT_{t-4} - 0.095\ S_{t-8}$$
$$(0.002)\qquad (0.039)$$

France $R^2 = 0.39$, $Q(45) = 52.13(0.216)$

$$\Delta S_t = -0.237\ \Delta S_{t-1} + 0.021\ \Delta LONG_{t-2} - 0.028\ \Delta LONG^*_{t-1} - 0.030\ \Delta LONG^*_{t-5}$$
$$(0.089)\qquad (0.011)\qquad (0.012)\qquad (0.015)$$
$$+0.003\ \Delta TB_{t-3} + 0.004\ \Delta TB_{t-4} - 0.079\ S_{t-8} - 0.018\ LONG^*_{t-8}$$
$$(0.001)\qquad (0.002)\qquad (0.041)\qquad (0.009)$$

Germany $R_2 = 0.42$, $Q(45) = 43.24(0.546)$

$$\Delta S_t = -0.191\ \Delta S_{t-3} - 0.226\ \Delta S_{t-4} - 0.224\ \Delta S_{t-6} - 0.024\ \Delta LONG_{t-1} - 0.092\ S_{t-8}$$
$$(0.091)\qquad (0.094)\qquad (0.091)\qquad (0.012)\qquad (0.041)$$

Japan $R^2 = 0.40$, $Q(45) = 40.12(0.678)$

$$\Delta S_t = -0.184\ \Delta S_{t-7} + 0.012\ \Delta SHORT^*_{t-1} - 0.027\ \Delta LONG_{t-7} - 0.041\ \Delta LONG^*_{t-1}$$
$$(0.097)\qquad (0.006)\qquad (0.013)\qquad (0.012)$$
$$+0.025\ LONG^*_{t-7} - 0.151\ \Delta S_{t-8} - 0.013\ \Delta LONG^*_{t-8}$$
$$(0.012)\qquad (0.059)\qquad (0.006)$$

U.K. $R^2 = 0.45$, $Q(45) = 45.60(0.446)$

$$\Delta S_t = -1.666 + 0.481\ \Delta IP_{t-7} + 0.011\ \Delta SHORT^*_{t-1} + 0.001\ \Delta TB_{t-2}$$
$$(0.792)\ (0.214)\qquad (0.004)\qquad (0.0003)$$
$$-0.073 S_{t-8} + 0.514\ IP_{t-8} + 0.0003\ TB_{t-8}$$
$$(0.033)\qquad (0.189)\qquad (0.0001)$$

Notes: 1. S denotes spot exchange rate, M money supply, IP industrial production, SHORT short-term interest rate, LONG long-term interest rate and TB trade balance.

2. The asterisk indicates a foreign variable.

3. R^2 is the coefficient of determination; Q(45) is the Ljung-Box test statistic for serial correlation distributed as χ^2 with 45 degrees of freedom in our case; figures in parentheses beside Ljung-Box test statistics are marginal significance levels; those below coefficient estimates are standard errors.

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