Expectations of Exchange Rates and Differential Inflation Rates: Further Evidence on Purchasing Power Parity in Efficient Markets

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Expectations of Exchange Rates and Differential Inflation Rates: Further Evidence on Purchasing Power Parity in Efficient Markets

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ABSTRACT

The paper tests the null hypothesis of ex ante purchasing power parity. The empirical evidence obtained is inconsistent with the null for major industrialized countries over the current floating exchange rate regime. Expected nominal exchange rate changes appear to deviate systematically from expected inflation rate differentials over the same holding period even though real exchange rate changes appear to be serially uncorrelated. This supports the presence of time-varying risk premia in foreign exchange markets and real determinants of exchange rate movements as suggested by equilibrium theories of international asset markets.

THIS PAPER TESTS THE hypothesis that the expected change in the exchange rate of two countries equals the expected differential in their inflation rates over the same holding period. This relationship is an ex ante version of the classical (relative) purchasing power parity (PPP) that holds under some versions of efficient foreign exchange and international commodity markets. It has important implications for the efficacy of government policies in open economies and has been the focus of numerous empirical studies.

The ex ante PPP (EPPP) differs from the traditional PPP proposed by Cassel [4] in the explicit treatment of uncertainty. As formulated by Roll [32], the EPPP is based on the constraint that, in efficient markets, the net return to speculators engaging in intertemporal speculation on goods in the foreign country is anticipated to be zero.[1] An immediate implication of EPPP is that innovations in real exchange rates or deviations from PPP follow a martingale process. In sharp contrast to the well-documented empirical rejections of the conventional PPP, strong support for the martingale property of EPPP exists in the literature.[2]

The proposition that the differences in the expected price relatives of two countries are exactly accounted for by the expected change in exchange rates over the same period is inconsistent with a number of theoretical observations in the extant literature. In particular, the EPPP model assumes that the expected real interest rate differential between each pair of countries is intertemporally constant in order to rule out commodity and financial arbitrage in a world of

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[1] In contrast to Roll's cross-border commodity speculation, Adler and Lehmann [1] derived the ex ante model as a consequence of cross-border financial speculation in bonds.

[2] For empirical evidence on PPP, see, for example, Krugman [24], Dornbusch [10, 11], and Frenkel [14]. Empirical evidence on EPPP can be found in the work of Stockman [34], Roll [32], Darby [8], Adler and Lehmann [1], and others.

uncertainty (Roll [32], Adler and Lehmann [1]).[3] In constrast, numerous international intertemporal asset-pricing models have demonstrated that, when traders are risk averse, the same time-varying risk premium that separates the forward rate from the expected future spot rate also separates the expected real returns on default-free nominal bonds of two countries.[4] In addition, work with closed-economy models of rational expectations and completely flexible prices has shown that, in general, the assumption of constant expected real returns will be violated (for example, LeRoy [25] and Lucas [27]).[5] These results have been extended to an international context by Lucas [28]. Furthermore, Samuelson [33] has observed that the random-walk behavior of real stock prices with no upper bounds makes little sense if the real resources of the firm are finite. Similarly, if the value of one country's output in terms of another country's output is limited, real exchange rate movements are unlikely to be a martingale. In other words, real exchange rates may be determined primarily by real factors such as technological constraints and adjustment costs, which are unaccounted for by the martingale model.

Empirically, evidence contrary to the EPPP model also has been reported in the literature. Mishkin [31] obtained evidence unfavorable to the EPPP hypothesis when the model was examined in conjunction with other equilibrium conditions. In a recent paper, Cumby and Obstfeld [6] rejected strongly the proposition that expected inflation rate differentials and expected exchange rate changes are equal to one another in a direct test of the hypothesis. However, their analysis is confined to just five exchange rates in terms of the U.S. dollar from September 1975 to May 1981. In addition, Hodrick [21], Mishkin [31], and Cumby and Obstfeld [6] found evidence against the hypothesis that expected real interest rates are equal.[6]

The purpose of this paper is to examine the equality restriction between expected nominal exchange rate appreciations and expected inflation rate differentials imposed by the EPPP in light of the conflicting results previously reported. The next section formally states the EPPP model and highlights its testable implications. Section II contains a description of the data set used. Section III reports the results of tests of the martingale hypothesis. Section IV provides some evidence on the presence of conditional heteroskedasticity and extends the tests conducted by Cumby and Obstfeld [6] to additional exchange rates and longer time series using heteroskedasticity-consistent covariance matrix estimators. Section V ends the paper with a conclusion.

I. The Model

Roll [32] developed the concept of EPPP by invoking the assumption of efficient commodity markets. He considered the case of an investor in the home country

[3] Roll's version can also be criticized for requiring the speculator to hold the price indices.

[4] See, among others, the papers by Hodrick [22], Stulz [35], and Lucas [28] for international asset-pricing models with time-varying risk premia.

[5] Specifically, for the assumption to be valid, one has to rule out nonconstant returns to scale in production (as shown by LeRoy [25]) or discounting of the future (as shown by Lucas [27]), neither of which is attractive.

[6] See also the papers cited in footnote 9 for indirect evidence against the EPPP model.

who speculates by buying the foreign commodity at time t and selling it at time $t + 1$ in the foreign country. Therefore, the real return to an investor from intertemporal speculation (with no storage or shipping costs) is the percentage of change in the real exchange rate

$$r_{t+1} = Ds_{t+1} - (I_{t+1} - I^*_{t+1}),$$

where

$Ds_{t+1} = \ln(S_{t+1}) - \ln(S_t)$ = percentage of change in nominal exchange rates,
S_{t+1} = home currency price of foreign currency at time $t + 1$,
I_{t+1} = home inflation rate over the period t to $t + 1$, and
I^*_{t+1} = foreign inflation rate over the period t to $t + 1$.

The EPPP stipulates that all available information is utilized by the market participants so that the real return would be expected to be zero:

$$E[r_{t+1} \mid Q_t] = 0$$

or

$$E[I_{t+1} - I^*_{t+1} \mid Q_t] = E[Ds_{t+1} \mid Q_t], \tag{1}$$

where Q_t is the information set available at time t. In constrast to (1), the conventional PPP states that $r_{t+1} = 0$.

An immediate testable implication of (1) is that real exchange rates are mutually uncorrelated at all lags. This follows as $[r_{t-i}\colon i \geq 0]$ is included in Q_t.

An alternative testable implication of the martingale model involves expressing (1) in a regression format as

$$I_{t+1} - I^*_{t+1} = a + bE[Ds_{t+1} \mid Q_t] + u_{t+1}, \tag{2}$$

where $u_{t+1} = (I_{t+1} - I^*_{t+1}) - E[I_{t+1} - I^*_{t+1} \mid Q_t]$ is the innovation in inflation rate differentials. Therefore, the null hypothesis is $(a\ b) = (0\ 1)$.

The EPPP (1) requires the difference between real interest rates to be a constant. This has direct implication on the presence of a risk premium in the foreign exchange markets. A well-known riskless arbitrage condition in international markets is the interest rate parity relationship.[7] It equates forward premiums to nominal interest rate differentials of similar assets denominated in different currencies and can be combined with (1) to show that, under the constancy assumption for ex ante real rate differentials, the divergence of forward rates from expected future spot rates is a constant. However, modern asset-pricing models of exchange rate determination have shown that the forward rates are unbiased predictors of expected future spot rates only under risk neutrality and that a time-varying risk premium is a direct implication of risk aversion.[8] Recent empirical studies also support the hypothesis that a time-varying risk premium separates the forward rates from expected future spot rates.[9] Thus, the

[7] Marston [29], Frenkel and Levich [15, 16], and others have documented empirical evidence consistent with interest rate parity.

[8] Some of this literature is cited in footnote 4.

[9] See, for example, the work of Hansen and Hodrick [19], Cumby and Obstfeld [4], Hakkio [17], and Longworth [26].

results of this paper provide indirect evidence on the existence of time-varying risk premia in the foreign exchange markets.

II. The Data

The data of eleven major industrialized countries are used: the United States, Belgium, Canada, Denmark, France, Germany, Italy, Japan, the Netherlands, Switzerland, and the United Kingdom. This resulted in a total coverage of fifty-five exchange rates in the study. They are obtained from the International Financial Statistics tapes. Nominal exchange rates are end-of-month data, and consumer price indices are used to construct inflation rates. The starting date of the data is April 1971, which corresponds to Roll's [32] starting date for the flexible exchange rate period. The ending date is December 1984.

III. Tests of Real Exchange Rate Behavior

The EPPP stipulates that the innovations in real exchange rates are serially uncorrelated. This section tests the implied restriction for the present data set.

The Box-Pierce [3] Q-statistic is used to test the null hypothesis that the first n sample autocorrelations of real exchange rate changes are jointly insignificantly different from zero. The Q-statistic is computed as

$$Q(n) = T \sum_{i=1}^{n} \hat{\rho}_i^2,$$

where $\hat{\rho}_i = \sum r_t r_{t-i} / \sum r_t^2$ and T is the sample size. Under the null hypothesis, the statistic is distributed asymptotically as χ^2 with n degrees of freedom.

The Q-statistics are computed for $n = 6$, 12, 18, and 24. They all give similar conclusions, and only the results for 12 lags are presented in Table I. The marginal significance levels or p-values indicate that the null cannot be rejected for the vast majority of the cases at the usual significance levels. These results are consistent with those found by Stockman [34], Roll [32], Darby [8], Adler and Lehmann [1], and others.

IV. Tests of Nominal Exchange Rate and Inflation Rate Behavior

This section tests the H_0: $(a \; b) = (0 \; 1)$ imposed by EPPP. This cannot be accomplished by estimating (2) directly since data on $E[Ds_{t+1} \mid Q_t]$ are unavailable. However, $E[Ds_{t+1} \mid Q_t]$ can always be dichotomized as

$$E[Ds_{t+1} \mid Q_t] = Ds_{t+1} - v_{t+1}, \tag{3}$$

where v_{t+1} is the innovation in nominal exchange rate changes. Therefore, the unobservable variable can be replaced by its realization to obtain

$$I_{t+1} - I_{t+1}^{*} = a + bDs_{t+1} + e_{t+1}, \tag{4}$$

where $e_{t+1} = u_{t+1} - bv_{t+1}$. Under the null, e_{t+1} is the expectational error for real exchange rates.

Because e_{t+1} is correlated with Ds_{t+1}, using OLS to estimate (4) will lead to biased and inconsistent estimates. A procedure to obtain consistent estimates is

Table I

Q-Test for Serial Correlation of Real Exchange Rate Changes

Countries	Statistic	p-Value	Countries	Statistic	p-Value	Countries	Statistic	p-Value
BE/US	9.61	0.65	DE/CA	9.20	0.69	NE/FR	18.43	0.10
CA/US	29.85	0.00	FR/CA	8.56	0.74	SW/FR	5.26	0.95
DE/US	8.12	0.78	GE/CA	8.94	0.71	UK/FR	9.67	0.65
FR/US	8.12	0.78	IT/CA	5.94	0.92			
GE/US	7.31	0.84	JA/CA	10.09	0.61	IT/GE	6.33	0.90
IT/US	3.34	0.99	NE/CA	8.68	0.73	JA/GE	7.93	0.79
JA/US	13.05	0.37	SW/CA	5.02	0.96	NE/GE	24.88	0.02
NE/US	8.61	0.74	UK/CA	18.50	0.10	SW/GE	11.05	0.53
SW/US	4.57	0.97				UK/GE	9.52	0.66
UK/US	9.72	0.64	FR/DE	6.68	0.88			
			GE/DE	9.80	0.63	JA/IT	8.24	0.77
CA/BE	11.29	0.50	IT/DE	8.02	0.78	NE/IT	12.69	0.39
DE/BE	13.52	0.33	JA/DE	8.89	0.71	SW/IT	6.22	0.90
FR/BE	36.22	0.00	NE/DE	29.82	0.00	UK/IT	15.22	0.23
GE/BE	25.39	0.01	SW/DE	7.67	0.81	NE/JA	11.82	0.46
IT/BE	10.04	0.61	UK/DE	6.66	0.88	SW/JA	7.74	0.81
JA/BE	9.49	0.66				UK/JA	10.25	0.59
NE/BE	12.27	0.42	GE/FR	30.80	0.00	SW/NE	10.22	0.60
SW/BE	12.67	0.39	IT/FR	17.86	0.12	UK/NE	11.00	0.53
UK/BE	8.78	0.72	JA/FR	12.46	0.41	UK/SW	8.16	0.77

Notes: The results are obtained for the period April 1971 to December 1984 using IMF monthly data for nominal exchange rates and consumer price indices. The test statistic is distributed asymptotically as χ^2 (12). BE, Belgium; US, U.S.; DE, Denmark; CA, Canada; NE, Netherlands; FR, France; IT, Italy; GE, Germany; UK, U.K.; SW, Switzerland; JA, Japan.

instrumental variable (IV) estimation as suggested by McCallum [30]. Equation (3) shows that any member of the information set Q_t can be used as an instrument to yield consistent estimates as it is uncorrelated with e_{t+1}. To express this more succinctly, stack the T observations on the variables and rewrite (4) as

$$y = X\beta + e, \tag{5}$$

where $y_{t+1} = I_{t+1} - I^*_{t+1}$, $\beta = (a\ b)'$, and $X_t = (1\ Ds_{t+1})$. Also, define Z_t to be a row vector of instrumental variables.

The standard IV estimation assumes that the composite residuals in the vector e are serially uncorrelated and homoskedastic. However, unlike the typical IV estimation, the instruments used are predetermined but not exogenous. Recent econometric work on time-series estimation of rational expectations models has highlighted the complications due to serial correlation and heteroskedasticity. Specifically, they result in inefficient parameter estimates and biased variance-covariance matrices so that standard inference procedures are invalid. Flood and Garber [13], Hansen and Sargent [20], and others have noted that the standard serial correlation corrections are inappropriate as they lead to inconsistent estimates, and Engle [12] and Hansen [18] have emphasized the problem of conditional heteroskedasticity. Therefore, it is important to check for these conditions.

The results of the previous section indicate that, under the tentative null hypothesis that $\beta = (0\ 1)'$, e_{t+1} is serially uncorrelated. However, in foreign exchange markets, turbulent times tend to cluster temporally, followed by sustained periods of relative calm. Therefore, conditional heteroskedasticity is likely to occur. Evidence of conditional heteroskedasticity in the innovations from exchange rate regression models has been found by Cumby and Obstfeld [6], Hodrick and Srivastava [23], and Domowitz and Hakkio [9].

A test under the null hypothesis could also be conducted for conditional homoskedasticity.[10] Conditional homoskedasticity implies that

$$E[e_{t+1}^2 \mid Q_t] = c, \tag{6}$$

where c is a constant. Equation (6) specifies that a regression of e_{t+1}^2 on any member of Q_t should yield a coefficient insignificantly different from zero. A natural choice of regressors is the lagged values of e_{t+1}. In Table II, I report the results from estimating the equation

$$e_{t+1}^2 = \alpha_1 + \alpha_2 e_t + \alpha_3 e_t^2 + \xi_{t+1}. \tag{7}$$

The F-statistic is used to test the hypothesis that $\alpha_2 = \alpha_3 = 0$. The results show that of the fifty-five exchange rates examined, conditional heteroskedasticity is present in twenty-four percent of the cases at the ten percent significance level. Of course, for those exchange rates that fail to reject, the test is a necessary but not a sufficient condition for homoskedasticity, and further experiments can be performed. However, the results obtained in Table II are sufficient to indicate that heteroskedasticity is potentially a serious problem so that, while coefficient estimates are consistent, standard inference procedures may be faulty.

The two-step two-stage least-squares (2S2SLS) estimation procedure developed by Cumby, Huizinga, and Obstfeld [7] and used by Cumby and Obstfeld [6] to estimate (5) can be utilized to provide correct hypothesis testing under very general forms of heteroskedasticity. The 2S2SLS estimate of β is

$$\hat{\beta} = (X'Z\hat{\Omega}^{-1}Z'X)^{-1}X'Z\hat{\Omega}^{-1}Z'y, \tag{8}$$

where $\hat{\Omega}$ is a consistent estimate of

$$\Omega = \lim_{T\to\infty}(1/T)E(Z'ee'Z). \tag{9}$$

Under standard regularity conditions, it can be shown that

$$\sqrt{T}(\hat{\beta} - \beta) \sim N[0, \operatorname{plim} T^2(Z'X)^{-1}\Omega(X'Z)^{-1}].$$

The first step in the 2S2SLS estimation technique involves obtaining a consistent estimate of Ω. Hansen [18] proves that a consistent estimate of Ω is

$$\hat{\Omega} = (1/T) \textstyle\sum_{k=-L}^{L} \sum_{t=1}^{T} Z_t'\hat{e}_t\hat{e}_{t-k}'Z_{t-k}, \tag{10}$$

where L is the assumed order for a moving-average process of $Z_t e_t$, and the residual $\hat{e}_t$ is obtained from estimating (5) by 2SLS.[11] It is important to emphasize

[10] A similar test was performed by Cumby and Obstfeld [6] in a different context.

[11] The results reported in the paper are for $L = 0$, as similar results are obtained for nonzero values of L. However, when L is set to some nonzero value, (10) does not necessarily need to be positive definite. In order to obtain a positive-definite matrix, the covariances for the nonzero lags are damped with the lag window $[(L + 1 - |k|)/(L + 1)]^q$. In general, a value of q less than one is adequate to

Table II

A Test for Conditional Heteroskedasticity of Real Exchange Rate Changes

Countries	Statistic	p-Value	Countries	Statistic	p-Value	Countries	Statistic	p-Value
BE/US	3.104	0.05	DE/CA	4.058	0.02	NE/FR	2.219	0.11
CA/US	1.490	0.23	FR/CA	1.239	0.29	SW/FR	3.059	0.05
DE/US	3.282	0.04	GE/CA	1.585	0.21	UK/FR	0.587	0.56
FR/US	0.911	0.40	IT/CA	0.217	0.80			
GE/US	2.031	0.13	JA/CA	0.565	0.57	IT/GE	2.594	0.08
IT/US	1.034	0.36	NE/CA	5.475	0.01	JA/GE	1.096	0.34
JA/US	1.608	0.20	SW/CA	1.601	0.20	NE/GE	4.241	0.02
NE/US	5.010	0.01	UK/CA	0.197	0.82	SW/GE	5.209	0.01
SW/US	1.951	0.15				UK/GE	0.877	0.42
UK/US	1.042	0.36	FR/DE	0.073	0.93			
			GE/DE	0.631	0.53	JA/IT	0.566	0.57
CA/BE	5.164	0.01	IT/DE	1.205	0.30	NE/IT	2.509	0.08
DE/BE	0.246	0.78	JA/DE	1.714	0.18	SW/IT	2.337	0.09
FR/BE	0.367	0.69	NE/DE	0.271	0.76	UK/IT	0.360	0.70
GE/BE	1.300	0.28	SW/DE	7.528	0.00	NE/JA	0.400	0.67
IT/BE	1.046	0.35	UK/DE	0.140	0.87	SW/JA	2.225	0.11
JA/BE	0.590	0.56				UK/JA	0.375	0.69
NE/BE	0.682	0.51	GE/FR	0.735	0.48	SW/NE	1.295	0.28
SW/BE	1.428	0.24	IT/FR	1.137	0.32	UK/NE	0.466	0.63
UK/BE	0.415	0.66	JA/FR	1.356	0.26	UK/SW	0.071	0.93

Notes: The results are obtained for the period April 1971 to December 1984 using IMF monthly data for nominal exchange rates and consumer price indices. The test statistic is the F-statistic. BE, Belgium; US, U.S.; DE, Denmark; CA, Canada; NE, Netherlands; FR, France; IT, Italy; GE, Germany; UK, U.K.; SW, Switzerland; JA, Japan.

that the procedure is valid for very general forms of heteroskedasticity and therefore does not require ad hoc heteroskedasticity specifications.[12]

In Table III, the results obtained using 2S2SLS are presented. The instruments used are one-period lagged inflation rate differentials between all industrialized countries and the reference currency. The null hypothesis $\beta = (0\ 1)'$ is examined using a chi-square statistic with two degrees of freedom. When the denominator currency of the exchange rates is the U.S. dollar, the null hypothesis is strongly rejected, with the exceptions of the Danish and Swiss exchange rates. The results also indicate that, when other currencies are used as the reference currency, the majority of the cases are rejected. Even for those cases that fail to reject the null hypothesis, all the slope coefficients are insignificantly different from zero and occasionally have the wrong sign.

The eighties were characterized by high real exchange rates in the major industrialized countries (see, for example, Blanchard and Summers [2]). Therefore, in order to determine whether the results are affected by the events of this

obtain a positive-definite matrix. Regression Analysis of Time Series econometric software is utilized for computing these estimates. Alternatively, as shown by Hansen [18], frequency-domain techniques can be used to estimate Ω. Specifically, a consistent estimate of Ω is provided by $2\pi\hat{s}(0)$, where $\hat{s}(0)$ is the estimate of the following spectral-density matrix evaluated at frequency zero:

$$\hat{s}(\omega) = 1/(2\pi)\ \textstyle\sum_{j=-\infty}^{+\infty} \exp(-i\omega j)(Z_t'\hat{e}_t\hat{e}_{t-j}'Z_{t-j}).$$

[12] It also is valid for very general forms of serial correlation in the disturbances.

Table III

Tests of Ex Ante PPP

	Intercept		Slope		Test	
Countries	Coefficient	Standard Error	Coefficient	Standard Error	Statistic	p-Value
BE/US	0.0002	0.0005	−0.0019	0.2673	24.05	0.00
CA/US	0.0006	0.0022	0.0967	1.3493	7.34	0.03
DE/US	0.0023	0.0028	−0.2262	1.0419	2.97	0.23
FR/US	0.0020	0.0016	0.0352	0.4821	19.18	0.00
GE/US	−0.0021	0.0004	−0.0067	0.3345	24.83	0.00
IT/US	0.0036	0.0082	0.2576	1.2081	5.30	0.07
JA/US	−0.0000	0.0028	0.0693	1.1562	9.51	0.01
NE/US	−0.0008	0.0005	−0.1385	0.5619	8.11	0.02
SW/US	−0.0027	0.0063	−0.1881	2.2345	1.20	0.55
UK/US	0.0030	0.0018	0.0662	0.3718	10.58	0.01
CA/BE	0.0006	0.0004	−0.0182	0.2961	12.41	0.00
DE/BE	0.0012	0.0016	0.3001	1.2945	0.85	0.65
FR/BE	0.0018	0.0014	0.0991	0.6802	1.76	0.42
GE/BE	−0.0026	0.0017	−0.1606	0.7467	2.44	0.30
IT/BE	0.0038	0.0038	0.2720	0.7063	1.09	0.58
JA/BE	0.0005	0.0016	0.2303	0.3681	19.87	0.00
NE/BE	−0.0010	0.0013	−0.0222	0.8396	3.64	0.16
SW/BE	−0.0026	0.0024	−0.0862	0.5666	22.92	0.00
UK/BE	0.0029	0.0009	0.0773	0.2260	16.77	0.00
DE/CA	0.0011	0.0012	−0.2151	0.9865	1.52	0.47
FR/CA	0.0013	0.0007	0.0236	0.4019	6.23	0.04
GE/CA	−0.0028	0.0013	0.1778	0.4895	4.75	0.09
IT/CA	0.0035	0.0066	0.2177	1.2466	1.18	0.55
JA/CA	0.0002	0.0081	0.3057	2.0731	8.15	0.02
NE/CA	−0.0017	0.0009	−0.1162	0.5280	4.48	0.11
SW/CA	−0.0034	0.0069	−0.1183	1.5408	8.26	0.02
UK/CA	0.0026	0.0013	−0.0127	0.3369	10.13	0.01
FR/DE	0.0005	0.0010	−0.1569	1.1812	1.36	0.51
GE/DE	−0.0034	0.0026	0.1046	0.7743	2.07	0.35
IT/DE	0.0031	0.0025	0.1380	0.6219	2.25	0.32
JA/DE	−0.0012	0.0019	0.1464	0.3379	21.07	0.00
NE/DE	−0.0031	0.0022	−0.2341	0.8161	2.35	0.31
SW/DE	−0.0048	0.0075	−0.1827	1.3866	6.18	0.05
UK/DE	0.0015	0.0008	0.0432	0.2990	10.40	0.01
GE/FR	−0.0038	0.0015	0.1001	0.3606	6.25	0.04
IT/FR	0.0023	0.0043	0.2905	1.2039	0.52	0.77
JA/FR	−0.0010	0.0026	0.2224	0.4267	21.10	0.00
NE/FR	−0.0032	0.0014	−0.1048	0.4127	8.30	0.02
SW/FR	−0.0039	0.0068	0.0499	1.0941	20.31	0.00
UK/FR	0.0011	0.0006	0.1369	0.2924	9.13	0.01
IT/GE	0.0056	0.0049	0.2408	0.6565	1.34	0.51
JA/GE	0.0024	0.0018	0.3471	0.9714	11.81	0.00
NE/GE	0.0015	0.0011	−0.3319	1.2943	2.45	0.29
SW/GE	−0.0002	0.0011	−0.0529	0.5211	35.07	0.00
UK/GE	0.0055	0.0012	−0.0161	0.2094	23.55	0.00
JA/IT	−0.0037	0.0053	0.2068	0.5856	17.00	0.00
NE/IT	−0.0050	0.0037	0.1737	0.5292	3.68	0.16
SW/IT	−0.0053	0.0100	0.2335	1.0242	10.68	0.00
UK/IT	−0.0015	0.0020	0.2372	0.7241	1.24	0.54
NE/JA	−0.0009	0.0010	0.1587	0.3087	26.79	0.00
SW/JA	−0.0017	0.0012	0.3150	0.9390	1.98	0.37

Table III (cont.)

Tests of Ex Ante PPP

	Intercept		Slope		Test	
Countries	Coefficient	Standard Error	Coefficient	Standard Error	Statistic	p-Value
UK/JA	0.0035	0.0021	0.0056	0.2791	30.63	0.00
SW/NE	−0.0019	0.0022	−0.2122	0.7516	9.33	0.01
UK/NE	0.0043	0.0012	−0.0381	0.2034	29.68	0.00
UK/SW	0.0057	0.0020	−0.0387	0.2563	24.19	0.00

Notes: The results are obtained for the period April 1971 to December 1984 using IMF monthly data for nominal exchange rates and consumer price indices. The test statistic is distributed asymptotically as χ^2 (2). BE, Belgium; US, U.S.; DE, Denmark; CA, Canada; NE, Netherlands; FR, France; IT, Italy; GE, Germany; UK, U.K.; SW, Switzerland; JA, Japan.

period, the tests were repeated for the period April 1971 to September 1979. The results obtained reject the null hypothesis more strongly than for the April 1971 to December 1984 sample. For example, all the U.S. foreign exchange rates are rejected, and only those of Danish krone and Dutch guilder fail to reject for the Belgium exchange rates.[13] Thus, the results cast doubt on the validity of EPPP and support the findings of Cumby and Obstfeld [6].

V. Conclusion

This paper has investigated the null hypothesis of expected nominal exchange rate changes as unbiased predictors of expected inflation rate differentials over the same horizon. On the whole, the empirical evidence is inconsistent with the null for major industrialized countries over the current floating exchange rate regime. This supports the presence of time-varying risk premia in foreign exchange markets and real determinants of exchange rate movements as suggested by equilibrium theories of international asset markets. Of course, since the tests performed are only meaningful when examined in conjunction with market efficiency, the results may simply indicate a failure of rational expectations.

[13] These results are available from the author upon request.

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