E C O N O M I C S B U L L E T I N

Efficient unit root tests of real exchange rates in the post−Bretton Woods era

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Abstract

We apply the efficient unit root tests of Elliott, Rothenberg, and Stock (1996), and Elliott (1999) to twenty−one real exchange rates using monthly data of the G−7 countries from the post−Bretton Woods floating exchange rate period. Our results indicate that, for eighteen out of the twenty−one real exchange rates, the null hypothesis of a unit root can be rejected at the 10% significance level or better using the the Elliott et al. (1996) test. Using the Elliott (1999) test, we have only nine rejections out of the twenty−one real exchange rates at the 10% significance level or better. We also find no strong evidence to suggest that the use of non−U.S. dollar based real exchange rates tend to produce more favorable result for long−run PPP than the use of U.S. dollar based real exchange rates as Lothian (1998) has concluded.

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1. Introduction

The theory of purchasing power parity (PPP) has a long tradition in international economics. It is a central building block in the monetary models of exchange rate determination. In the monetary approach, e.g., Frenkel (1978), PPP is assumed to hold continuously. This assumption, however, is at odd with the frequently observed deviations of nominal exchange rates from their implied PPP levels. In Dornbusch's (1976) sticky-price monetary model, because prices are sticky in the short run, PPP is assumed not to hold. For the long run, however, PPP is still a maintained assumption. Dornbusch's (1976) model can provide an explanation for the short-run volatility of real exchange rates, but the estimated half-life of three to five years for the shocks to damp out appears to be more consistent with real rather than nominal shocks.[1] Given the central place that PPP plays in the monetary models of exchange rate determination, it is not surprising that considerable research has been devoted to its empirical verification. Moreover, knowledge of the stochastic structure of real exchange rates is useful in determining whether shocks to real exchange rates are permanent or transitory, and whether they are more consistent with monetary or real shocks.

Recent empirical tests of PPP have mainly focused on the long run given that there are frequent large and persistence short-run deviations from PPP. Earlier empirical results on PPP have not been very encouraging, especially for the post-Bretton Woods period [see the survey by Rogoff (1996)]. More recent empirical results appear to be more encouraging, but they are not very robust, however. Two main explanations have been offered for why the earlier studies found mostly unfavorable evidence of long run PPP for the post-Bretton Woods period. First, Lothian (1998) noted that there was a substantial real appreciation of the U.S. dollar for 1980-1985, and an almost equal offsetting real depreciation for 1985-1987, leading him to conclude that the frequent failures to find evidence in favor of long-run PPP with post-Bretton Woods data is not a generic problem to this period. Rather, it is confined to using the U.S. dollar as the base currency and is restricted to the early to mid-1980s. Indeed, when German mark is used as the base currency, researchers tend to find more favorable results for long-run PPP [see Papell (1997), and Papell and Theodoridis (1998) for recent examples]. Second, the augmented Dickey-Fuller (ADF, 1981) test used in many of the earlier studies is known to have low power in small samples against plausible alternatives [see, for example, Hakkio (1986), and DeJong, Nankervis, Savin, and Whiteman (1992)]. Thus, it is not possible to distinguish between the low power of the test employed or that PPP does not hold in the post-Bretton Woods floating period.

One way to increase the power of the empirical tests is to use longer span of data, in some cases going back to the 1790s. For example, Diebold, Husted, and Rush (1991),[2] and Lothian and Taylor (1996), found evidence to support long-run PPP, Engel and Kim (1999), found evidence of a permanent (i.e., a unit root) component in the U.S./U.K. real exchange rate. Rogoff (1996) and others have noted that the economic implications of these studies are unclear since they typically mix fixed and floating exchange rates data. Moreover, long spans of time series data may potentially contain serious structural breaks. Engel (1996) also argued that these

[1] A recent paper by Murray and Papell (2002) has shown that the half-life estimates are extremely unreliable, however.

[2] It should be noted that in addition to using long span of data, Diebold, Husted and Rush (1991) also used fractional differencing in their model. Thus, it is not possible to distinguish whether their finding of favorable evidence of long-run PPP is due to the long span of data or to the use of autoregressive fractionally integrated moving-average process (ARFIMA).

studies could have serious size biases, and may fail to reject a sizable unit root. Finally, these studies also do not shed much light on the question of whether or not PPP is a valid hypothesis in the post-Bretton Woods floating period.

Researchers have also used panel data to increase the power of unit root tests. Recent studies finding evidence to support long-run PPP using post-Bretton Woods data include studies by Jorion and Sweeney (1996), Papell (1997), Papell and Theodoridis (1998), and Koedijk, Schotman, and Van Dijk (1998). On the other hand, O'Connell (1998) found little evidence to support long-run PPP after accounting for serial correlation. Rogoff (1996) and Papell (1997) found that panel unit root tests results are sensitive to the size of the panel and the countries included. Karlsson and Löthgren (2000) also questioned the power of panel unit root tests. Using Monte Carlo simulations, they found that for panels with long spans of data, the null hypothesis of unit roots could be erroneously rejected even when only a small proportion of the series is stationary. For panels with short spans of data, they found that the null hypothesis is frequently not rejected even when a large fraction of the series is stationary. Thus, they concluded that the rejection or the non-rejection of the null hypothesis of unit roots do not provide sufficient evidence to conclude that all the series in the panels are stationary or that they all have a unit root. An exception, however, is the panel study by Sarno and Taylor (1998). Using a special application of the Johansen (1988) Likelihood ratio, where the null hypothesis of a unit root is rejected only when all the series are stationary, Sarno and Taylor (1998) concluded that their four real exchange rates are jointly stationary series. Thus, panel unit root tests have produced encouraging but inconclusive results.

Still other researchers, using different empirical methodologies, found mixed results for the post-Bretton Woods period. For example, Cheung and Lai (1998), Cushman (2001), and Amara and Papell (2003), using more efficient unit root tests, found more encouraging results than when using the ADF tests. Culver and Papell (1999), using the tests proposed by Kwiatkowski, Phillips, Schmidt and Shin (KPSS, 1992) where stationarity is the null, rather than the alternative hypothesis, found favorable evidence to support long-run PPP. Caner and Kilian (2000) found that the KPSS test could have serious distortion, however. Finally, Baum, Barkoulas, and Caglayan (1999), allowing for fractional differencing or structural breaks, found no evidence to support long-run PPP.

Our brief review suggests that recent empirical studies have tended to be more supportive of long-run PPP than earlier studies for the post-Bretton Woods period. However, the results are not very robust, and consistent individual country time series evidence from this period continues to be scarce.

We have two purposes in this study. First, we test for long-run PPP with data from the post-Bretton Woods period using the efficient unit root tests proposed by Elliott, Rothenberg, and Stock (1996), and Elliott (1999).[3] The efficient unit root tests provide substantial power gain over the Dickey-Fuller ADF test and thus provide an alternative to using long span of data or

[3] Our study differs from the studies by Cheung and Lai (1998), Cushman (2001), and Amara and Papell (2003) in several important respects. We use monthly data and extend Cheung and Lai's monthly data by four years, while Cushman, and Amara and Papell used quarterly data. All the authors, except Cushman, used several efficient unit root tests, including the DF-GLS test of Elliott *et al*. (1996), but none used the efficient unit root test of Elliott (1999). Thus, we provide a comparison of these two efficient unit root tests. We provide a more systematic assessment of Lothian's (1998) conclusion by using seven different base currencies and twenty-one real exchange rates. Cheung and Lai used four base currencies and ten real exchange rates, Cushman used two base currencies, while Amara and Papell used only U.S. dollar-based real exchange rates.

panel unit root tests, allowing us to avoid the problems associated with these empirical methodologies as discussed earlier. Surprisingly, however, there has been very little empirical work on unit root testing using efficient unit root tests. Second, we provide a more systematic assessment of Lothian's (1998) assertion that the failure of the earlier studies to find favorable evidence of long-run PPP in the post-Bretton Woods period is likely due to the use of U.S. dollar as the base currency.

In the next section, section 2, we discuss the efficient unit root tests. In section 3, we discuss our data set and present our empirical results. Summary and conclusions are in section 4.

2. Efficient unit root tests

We start by defining the real exchange rate in natural logarithm form as:

$$q_t = e_t + P_t^* - P_t, \tag{1}$$

where q_t is the natural logarithm of the real exchange rate, e_t is the natural logarithm of the nominal exchange rate, defined as the domestic currency price of one unit of foreign currency, P_t is the natural logarithm of an index of the domestic price level, and P_t^* is the natural logarithm of an index of the foreign price level. Conventional ADF tests of long-run PPP involve estimating an equation similar to

$$q_t = \alpha_0 + \rho q_{t-1} + \sum_{i=1}^{k} \alpha_i \Delta q_{t-i} + \varepsilon_t, \tag{2}$$

where α_0 is a constant, Δ is the first difference operator, i.e., $\Delta q_t = q_t - q_{t-1}$, and ε_t is a serially uncorrelated error process. Long-run PPP requires that $\hat{\rho} < 1$. If $\hat{\rho} = 1$, there is a unit root in the real exchange rate series, shocks to the real exchange rate are permanent and long-run PPP does not hold.

The efficient unit root tests of Elliott *et al.* (1996) and Elliott (1999) are similar, differing only in the initial condition assumption. The efficient unit root test of Elliott *et al*. (1996) is based on the point optimal tests. In general, while no uniformly most powerful unit root test of $H_O: \rho = 1$ against the general alternative $H_A: \rho < 1$ exists, there is an optimal test, however, against a specific local alternative $H_A: \rho = \bar{\rho} < 1$, where $\bar{\rho} = 1 + \bar{c}/T$, $\bar{c} < 0$ is a specific constant, and T is the sample size. Using a sequence of Neyman-Person tests of the null hypothesis of a unit root against a set of stationary local alternatives, Elliott *et al*. (1996) derived the asymptotic maximal power envelope. From the power calculations, Elliott *et al*. (1996) shown that substantial power gain over the standard ADF test could be obtained from a modified ADF test, which they called the DF-GLS test. The DF-GLS test involves estimating the following equation with ordinary least squares:

$$\Delta q_t^d = \delta_0 q_{t-1}^d + \sum_{i=1}^{k} \delta_i \Delta q_{t-i}^d + \varsigma_t^d, \tag{3}$$

where ς_t^d is a serially uncorrelated error process, q_t^d is the locally detrended series of q_t, where

$$q_t^d = q_t - z_t \beta, \tag{4}$$

and $z_t = (1, t)$, for the locally detrended series with a constant and a linear trend, and $z_t = 1$, for series without a linear trend. Finally, β is the vector of least squares regression coefficients of

$\tilde{q}_t = [q_1, (1-\bar{\rho}L)q_2, \ldots, (1-\bar{\rho}L)q_T]'$ on $\tilde{z}_t = [z_1, (1-\bar{\rho}L)z_2, \ldots, (1-\bar{\rho}L)z_T]'$, and L is the lag operator, i.e., $Lz_t = z_{t-1}$. A t-test is used to test the null hypothesis $H_0 : \delta_0 = 0$ against $H_A : \delta_0 < 1$.

Elliott's (1999) efficient unit root test, denoted as DF-GLS^{u}, differs from Elliott *et al.* (1996) in its assumption about the initial value of the alternative model. Specifically, both Elliott *et al.* (1996) and Elliott (1999) assume that their data $\{y_t\}_{t=1}^{T}$ are generated according to

$$y_t = d_t + u_t \text{, and} \tag{5}$$

$$u_t = \rho u_{t-1} + v_t, \tag{6}$$

where d_t is a deterministic component which may or may not contain a deterministic linear trend, and v_t is a stationary error process which may or may not be serially correlated. Elliott *et al.* (1996) assumed that the initial value of u_t, i.e., u_0 is zero both when $\rho = 1$ and when $\rho < 1$ in equation (6), so that $u_1 = v_1$. Elliott (1999) assumed that u_0 is zero when $\rho = 1$, so that $u_1 = v_1$ also, but when $\rho < 1$, u_1 has mean zero and variance $\dfrac{Var(v_t)}{(1-\rho^2)}$. Since the alternative assumption involves the unknown parameter ρ, Elliott (1999) has shown that, since this unknown parameter does not disappear asymptotically, the likelihood test statistics and the power of the tests will differ from the optimal test in Elliott *et al.* (1996), and a different set of the critical values of the test statistics are derived in Elliott (1999). To implement the DF-GLS^{u} test, equation (3) is estimated by least squares, with q_t^d, z_t, β are as defined before, except that now $\tilde{q}_t = [(1-\bar{\rho}^2)^{1/2} q_1, (1-\bar{\rho}L)q_2, \ldots, (1-\bar{\rho}L)q_T]'$, and $\tilde{z}_t = [(1-\bar{\rho}^2)^{1/2} z_1, (1-\bar{\rho}L)z_2, \ldots, (1-\bar{\rho}L)z_T]'$.

In practice, however, it is difficult to know whether the sample data on hand conform to the data generating models of Elliott *et al.* (1996) or Elliott (1999). This is certainly an interesting area for further research. We report test results for both in our paper. Following Elliott *et al.* (1996), we use $\bar{c} = -7$ for test with a constant, and $\bar{c} = -13.5$ for test with a constant and a linear trend. Asymptotic critical values of the test statistics are derived in Elliott *et al* (1996) for the DF-GLS test. For the DF-GLS^{u} test, we also follow Elliott (1999) and use $\bar{c} = -10$ in both test with a constant, and test with a constant and a linear trend, and use the asymptotic critical values of the test statistics reported in Elliott (1999).

3. Empirical Results

The source of our data is SourceOECD, and our sample consists of the OECD G-7 countries. Our data consist of monthly observations from April 1973 to February 1999 for the G-7 countries,[4] and are not seasonally adjusted. The G-7 countries are the U.S., the U.K., Canada, Germany, Italy, France, and Japan. In all cases, we use the consumer price index as our measure of the average price level. The bilateral nominal exchange rates are monthly averages,

[4] SourceOECD recently changed its reporting method for exchange rates. Countries belonging to the Euro zone will now have only the euro/U.S. dollar nominal exchange rate reported from January 1999 on. Prior to this change, it was still possible to obtain individual country's exchange rate with the U.S. dollar even after the establishment of the Euro zone. Our data are the longest that we can find that still report individual country's exchange rate with the U.S. dollar.

and all use the U.S. dollar as the base currency, i.e., foreign currency per U.S. dollar. To test Lothian's (1998) assertion, we have also computed real exchange rates based on the pound sterling, the Canadian dollar, the German mark, the Italian lira, the France franc, and the Japanese yen. These non-U.S. dollar based exchange rates are computed as cross-rates.[5]

We start our empirical tests by first presenting in Table 1 our test for unit root using the ADF test. These results will provide an update on previous research through the beginning of 1999, and it is interesting to find out whether an addition of several more years of data would have made a difference in the ADF tests for unit root. Moreover, the ADF unit root test results will also provide a comparison to the efficient unit root tests of Elliott *et al*. (1996), and Elliott (1999). Note also that we provide unit root test results for forty-two real exchange rates (six real exchange rates for each of the seven currencies). Of course, there are only twenty-one different real exchange rates since the real exchange rate of country A's currency per unit of country B's currency is simply the inverse of the real exchange rate of country B's currency per unit of country A's currency. This is done so that we can examine how real exchange rates based on non-U.S. dollar would behave compared to the U.S. dollar based real exchange rates.

The ADF regression actually estimated is

$$\Delta q_t = \beta_0 + \lambda q_{t-1} + \sum_{i=1}^{l} \beta_i \Delta q_{t-i} + \beta_{l+1} t + \eta_t , \tag{7}$$

where t = a linear deterministic time trend, and η_t is a serially uncorrelated error process with zero mean and constant variance. The lag length for the lagged first-differences is determined by using a general-to-specific method recommended by Ng and Perron (1995) and Perron (1997). We start by estimating Equation (7) with a pre-determined maximum lag length of 12 and sequentially drop the last included lag if it is not statistically significant at the 10% significance level. If, however, the lag length determined is the same as the maximum lag length, we start over with a maximum lag length of 14. Note that we have included a linear time trend to allow for the possibility that the real exchange rates may be trend-stationary.[6] We report our empirical results for both with and without a linear time trend, however.

Column 2 of Table 1 shows the lag length chosen for the models. In columns 3 and 4, we show the t-statistic for the hypothesis $H_0 : \lambda = 0$ without and with a linear time trend, respectively. We use the critical values at the 5% and the 10% significance levels from Fuller (1976) and the lag-adjusted critical values for exact sample size from Cheung and Lai (1995). For the ADF test results without a linear time trend, the null hypothesis is not rejected at the 5% significance level for all cases. At the 10% significance level, the null hypothesis is rejected for three of twenty-one real exchange rates. When the ADF regression includes a linear time trend, the null hypothesis is rejected for three of twenty-one real exchange rates at the 5% significance level, and an additional two real exchange rates at the 10% significance level. In sum, a total of

[5] This assumes cross-rate equality except for transaction costs. This is probably a valid assumption for the G-7 countries. Alternatively, as long as the measurement error is a stationary process, our tests for unit-root will not be affected.

[6] Some researchers, e.g., Cheung and Lai (1998), and Koedijk, Schotman, and Van Dijk (1998), have found that the stochastic processes of some of the real exchange rates cannot be adequately modeled without the inclusion of a linear deterministic time trend. The linear deterministic time trend is generally interpreted as representing systematic differences in productivity growth between tradable and non-tradable goods in the two countries. On the other hand, other researchers, e.g., Papell and Theodoridis (1998), and Amara and Papell (2002) consider a linear time trend in the real exchange rate as inconsistent with long-run PPP.

seven of twenty one, or 33% of the real exchange rates, are found to be either stationary or trend-stationary at the 10% significance level or better using the conventional ADF approach. These results are broadly consistent with those found by other researchers. Thus, the addition of a few more years of monthly data appears to have no significant impact on the power of the ADF unit root test.

There is some evidence to suggest that non-U.S. dollar-based real exchange rates, e.g., Japanese yen-based and German mark-based real exchange rates, do appear to provide more favorable long-run PPP results. The improvement over U.S. dollar-based real exchange rates is marginal at best, however. Moreover, using pound sterling-based or Italian lira-based real exchange rates produce equally dismal results as using U.S. dollar-based real exchange rates. Thus, our ADF results provide no support to Lothian's (1998) assertion that the frequent failures to find favorable evidence of long-run PPP in earlier studies for the post-Bretton Woods period is confined to using U.S. dollar as the base currency.

Table 2 reports the results for both the DF-GLS and the DF-GLS^u unit root tests from estimating equation (3). The lag lengths used in these two tests are the same as those determined for the ADF test reported in Table 1. What we report in Table 2 are the t-statistics for the null hypothesis $H_0 : \delta_0 = 0$ against the alternative hypothesis $H_A : \delta_0 < 0$. Starting with the DF-GLS test without trend, at the 5% significance level, the null hypothesis of a unit root is rejected in five cases. At the 10% significance level, the null hypothesis is rejected for three additional real exchange rates. When a linear trend is included, the null hypothesis is rejected at the 5% significance level for six real exchange rates, and the null hypothesis is rejected at the 10% significance level for nine additional real exchange rates. In all, the DF-GLS test suggests that eighteen of twenty-one real exchange rates, or about 86% of the cases, are either stationary or trend-stationary at the 10% significance level or better. This represents a significant improvement over the results of the conventional ADF test. The most favorable results come from the franc-based and the yen-based real exchange rates where in each case all six real exchange rates are either stationary or trend-stationary at the 10% significance level or better. This is followed by the U.S. dollar-based, the Canadian dollar-based, and the mark-based real exchange rates with five out of six real exchange rates each, and the lira-based real exchange rates with four out of six cases. Overall, we conclude that there are no major differences in results between U.S. dollar-based and non-U.S. dollar-based real exchange rates, with perhaps the Italian lira-based real exchange rates being the exceptions.

Our results are slightly better than Cheung and Lai's (1998) DF-GLS test results. They found eight out of ten real exchange rates that are either stationary or trend-stationary at the 10% significance level or better while all ten of Cheung and Lai's real exchange rates in our sample are either stationary or trend-stationary at the 10% significance level or better. A comparison with Cushman's (2001) DF-GLS test results is more difficult since he included real exchange rates that are not in our sample and we have real exchange rates that are not in his sample. Nevertheless Cushman's and our results appear to be consistent with each other's. A comparison with Amara and Papell's (2003) DF-GLS results is even more difficult since they did not allow for a linear time trend in their regressions and only considered U.S. dollar-based real exchange rates. If we focus only on Amara and Papell's and our DF-GLS results without linear time trend, they have ten of twenty cases of rejection of the null hypothesis, compared to our eight of twenty-one cases at the 10% significance level or better. Finally, our results also confirm the results of Cheung and Lai (1998), and Amara and Papell (2003), that the use of the more efficient DF-GLS tests provide much more favorable evidence of long-run PPP for the post-Bretton Woods period than the use of the ADF tests.

We next turn to the DF-GLS^u results in Table 2. First thing to note is that, compared to the DF-GLS results, there are fewer cases where the null hypothesis of a unit root is rejected. In all, only nine out of twenty-one real exchange rates, or about 43%, can be characterized as either stationary or trend-stationary at the 10% significance level or better. These nine real exchange rates include the seven that were found to be either stationary or trend-stationary using the ADF tests. More importantly, no additional real exchange rate is found to be stationary or trend-stationary that was not found by the DF-GLS test. These results are slightly better than the results from the ADF tests, but are clearly much worst than the results using the DF-GLS tests. Unfortunately, there is currently no practical way to determine whether the DF-GLS test or the DF-GLS^u is more appropriate for our data sample, and thus it is difficult to choose between these two sets of results. It may be the case that the DF-GLS test is a more powerful test than the DF-GLS^u test. This will be an interesting area for future research. However, since we are looking for evidence to support long-run PPP in the post-Bretton Woods era, we are inclined to accept the more favorable results provided by the DF-GLS test.

4. Summary and Conclusions

Recent empirical studies, using different empirical methodologies to overcome the weaknesses of the conventional ADF tests, have found increasingly supportive evidence of long-run PPP. Consistent individual time series evidence from the post-Bretton Woods era, however, continues to be scarce. We applied the efficient tests for a unit root proposed by Elliott *et al.* (1996) and Elliott (1999) to twenty-one time series of real exchange rates using monthly data from the post-Bretton Woods period of the G-7 countries. We find evidence to support long-run PPP at the 10% or better significance level for eighteen out of the twenty-one real exchange rates using the Elliott *et al.* (1996) DF-GLS test, but only nine out of twenty-one using the DF-GLS^u test of Elliott (1999).

Second, our results suggest that there are only marginal differences between U.S. dollar-based and non-U.S. dollar-based real exchange rates, both when the empirical results are supportive and not supportive of long-run PPP. Thus, we have no evidence to support Lothian's (1998) assertion that the failure to find favorable evidence of long-run PPP in earlier studies for the post-Bretton Woods period can be attributed to the use of U.S. dollar as the base currency. Rather, our results strongly suggest that, more than likely, it is due to the low power of the statistical tests employed.

Our study, even-though we find evidence favorable to long-run PPP, should be viewed as an exploratory study. One area that future research should address is the following. We have seen that the use of the DF-GLS test or the DF-GLS^u test can produce rather different results. At the moment, we know of no operational way to distinguish between when the DF-GLS or the DF-GLS^u test should be used. A resolution of this question should be helpful not only to researchers in international finance, but researchers in other areas as well.

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Table 1
Univariate ADF Test Results

	l	$t(\lambda)$ without trend	$t(\lambda)$ with trend
Base Currency: U.S. Dollar			
Canada	12	-1.373	-2.025
U.K.	11	-2.594**	-2.725
Germany	10	-2.331	-2.472
Italy	10	-2.507	-2.545
Japan	11	-1.932	-2.731
France	3	-2.249	-2.244
Base Currency: Canadian Dollar			
U.S.A.	12	-1.373	-2.025
U.K.	11	-2.014	-2.719
Germany	10	-2.642**	-2.660
Italy	1^{+}	-2.057	-3.298**
Japan	11	-1.779	-3.416**
France	5	-2.329	-2.663
Base Currency: Pound Sterling			
U.S.A.	11	-2.594**	-2.725
Canada	11	-2.014	-2.719
Germany	12	-2.196	-2.793
Italy	11	-2.356	-2.351
Japan	10	-2.035	-2.595
France	1	-2.273	-2.630
Base Currency: German Mark			
U.S.A.	10	-2.331	-2.472
Canada	10	-2.642**	-2.660
U.K.	12	-2.196	-2.793
Italy	2	-1.654	-1.917
Japan	8	-1.668	-3.598*
France	9	-3.066**	-4.499*
Base Currency: Italian Lira			
U.S.A.	10	-2.507	-2.545
Canada	1^{+}	-2.057	-3.298**
U.K.	11	-2.356	-2.351
Germany	2	-1.654	-1.917
Japan	3	-2.433	-3.101
France	3	-1.966	-1.991
Base Currency: Japanese Yen			
U.S.A.	11	-1.932	-2.731
Canada	11	-1.779	-3.416**
U.K.	10	-2.035	-2.595
Germany	8	-1.668	-3.598*
Italy	3	-2.433	-3.101
France	8	-1.983	-3.791*
Base Currency: French Franc			
U.S.A.	3	-2.249	-2.245
Canada	5	-2.329	-2.663
U.K.	1	-2.273	-2.630
Germany	9	-3.066**	-4.499*
Italy	3	-1.966	-1.991
Japan	8	-1.983	-3.791*

Note: *,** Denote the rejection of the null hypothesis at the 5% and the 10% significance levels, respectively.
$^{+}$ The lag length is 10 for the Canadian dollar/lira real exchange rate when estimated without a linear time trend.

Table 2
Efficient Unit Root Tests

	l	DF-GLS Without trend	DF-GLS With trend	DF-GLS^u Without trend	DF-GLS^u With trend
Base Currency: U.S. Dollar					
Canada	12	-0.152	-2.105	-1.555	-2.083
U.K.	11	-1.900**	-2.708**	-2.594**	-2.723
Germany	10	-2.183*	-2.473	-2.345	-2.482
Italy	10	-2.448*	-2.568**	-2.505**	-2.573
Japan	11	-0.890	-2.762**	-1.888	-2.752
France	3	-2.254*	-2.256	-2.263	-2.246
Base Currency: Canadian Dollar					
U.S.A.	12	-0.152	-2.105	-1.555	-2.083
U.K.	11	-0.685	-2.695**	-2.074	-2.690
Germany	10	-2.367*	-2.651**	-2.653**	-2.663
Italy	1^+	-1.349	-3.296*	-2.077	-3.316*
Japan	11	-0.430	-3.422*	-1.755	-3.424*
France	5	-1.855**	-2.678**	-2.341	-2.678
Base Currency: Pound Sterling					
U.S.A.	11	-1.900**	-2.708**	-2.594**	-2.723
Canada	11	-0.685	-2.695**	-2.074	-2.690
Germany	12	-1.175	-2.810**	-2.218	-2.794
Italy	11	-1.609	-2.164	-2.380	-2.275
Japan	10	-1.379	-2.619**	-2.016	-2.602
France	1	-1.525	-2.642**	-2.291	-2.643
Base Currency: German Mark					
U.S.A.	10	-2.183*	-2.473	-2.345	-2.482
Canada	10	-2.367*	-2.651**	-2.653**	-2.663
U.K.	12	-1.175	-2.810**	-2.218	-2.794
Italy	2	-1.215	-1.950	-1.668	-1.961
Japan	8	-0.218	-3.589*	-1.599	-3.574*
France	9	-2.234*	-4.115*	-3.077*	-4.218*
Base Currency: Italian Lira					
U.S.A.	10	-2.448*	-2.568**	-2.505**	-2.573
Canada	1^+	-1.349	-2.992*	-2.077	-3.048*
U.K.	11	-1.609	-2.164	-2.380	-2.275
Germany	2	-1.215	-1.950	-1.668	-1.961
Japan	3	-0.999	-3.007*	-2.398	-3.100*
France	3	-1.799**	-2.061	-1.972	-2.041
Base Currency: Japanese Yen					
U.S.A.	11	-0.890	-2.762**	-1.888	-2.752
Canada	11	-0.430	-3.422*	-1.755	-3.424*
U.K.	10	-1.379	-2.619**	-2.016	-2.602
Germany	8	-0.218	-3.589*	-1.599	-3.574*
Italy	3	-0.999	-3.007*	-2.398	-3.100*
France	8	-0.705	-3.786*	-1.938	-3.769*
Base Currency: French Franc					
U.S.A.	3	-2.254*	-2.256	-2.263	-2.246
Canada	5	-1.855**	-2.678**	-2.341	-2.678
U.K.	1	-1.525	-2.642**	-2.291	-2.643
Germany	9	-2.234*	-4.115*	-3.077*	-4.218*
Italy	3	-1.799**	-2.061	-1.972	-2.041
Japan	8	-0.705	-3.786*	-1.938	-3.769*

Note: See notes to Table 1.