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Efficient Unit Root Tests of real Exchange Rates in the Post-Bretton Woods Era

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Abstract

We apply the efficient unit-roots tests of Elliott, Rothenberg, and Stock (1996), and Elliott (1998) 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 10et al (1996) DF-GLS test. The unit-root null hypothesis is also rejected for one additional real exchange rate when we allow for one endogenously determined break in the time series of the real exchange rate as in Perron (1997). In all, we find favorable evidence to support long-run purchasing power parity in nineteen out of twenty-one real exchange rates. Second, we 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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Efficient Unit Root Tests of Real Exchange Rates in the Post-Bretton Woods Era

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 exchanges 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. 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. Yet another reason why it is important to know the stochastic property of real exchange rates has to do with what Rogoff (1996) has called the "purchasing power parity puzzle". The puzzle is that, while monetary shocks combined with sticky nominal prices or wages 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] Thus, a knowledge of the stochastic structure of real exchange rates is useful in determining whether shocks to real exchange rates are permanent or transitory, and if transitory, whether they are predominantly monetary or real shocks.

Until quite recently, empirical results on PPP have not been very encouraging, especially using data from 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. We have two purposes in this study. First, we test for long-run PPP using data from the post-Bretton Woods period using two statistical methods. We start with the efficient unit-root tests proposed by Elliott, Rothenberg, and Stock (1996), and Elliot (1998). Next we allow for a more flexible model structure by allowing for shifts in the mean or the trend, or both. Second, we provide a more systematic assessment to 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, we will briefly review some of the recent empirical studies on PPP. We discuss the efficient unit-root tests in section 3. In section 4, we discuss our data set and present our empirical results. In Section 5, we discuss and present our results using more flexible models that allow for structural breaks. Summary and conclusions are in section 6.

2. Review of Recent Empirical Results on Tests for Long-Run PPP

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. Two main explanations have been offered for why researchers failed to find favorable evidence of long run PPP in some of the earlier studies using data from the post-Bretton Woods period. First, it is known that when German mark is used as the base currency, researchers tend to find more favorable results for long-run PPP than when the U.S. dollar is the base currency [see Papell (1997), and Papell and Theodoridis (1998) for recent examples]. This leads Lothian (1998) to conclude that the frequent failures to find evidence in favor of long-run PPP in the post-Bretton Woods floating exchange rate period 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 when first there was a substantial real appreciation of the U.S. dollar for 1980-1985, and an almost equal offsetting real depreciation for 1985-1987. Second, the augmented Dickey-Fuller (ADF, 1981) test is one of the most generally used tests for PPP, but this test is known to have low power in small samples against plausible alternatives, especially against trend-stationary alternative [see, for example, Hakkio (1986), and DeJong, Nankervis, Savin, and Whiteman (1992)]. Thus, it is not possible to distinguish whether the failure to find PPP is due to the low power of the tests 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. For example, Diebold, Husted, and Rush (1991),[2] using data going back to the gold standard period, Lothian, and Taylor (1996), using data dating back to the1790s and early 1800s, found evidence to support long-run PPP. On the other hand, Engel and Kim (1999), using monthly data dating back to 1885, found evidence of a permanent (i.e., a unit-root) component in the U.S./U.K. real exchange rate. In addition, Rogoff (1996) and others have noted that studies that used long spans of data typically mix fixed and floating exchange rates data, and the economic implications of mixing data from the two exchange rate regimes are unclear. Moreover, long spans of time series data may potentially contain serious structural breaks. Engel (1996) also argued that these studies can 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.

Another way to increase the power of the unit-root tests is to use panel data. Recent examples using data from the post-Bretton Woods floating exchange rate period include studies by Jorion and Sweeney (1996), Papell (1997), Papell and Theodoridis (1998), and Koedijk, Schotman, and Van Dijk (1998). These studies all found evidence to support long-run PPP. On the other hand, O'Connell (1998) found little evidence to support long-run

PPP after accounting for serial correlation. Papell (1997) found that evidence in favor of long-run PPP is dependent on the size of the panel and the countries included. Rogoff (1996) also noted that with panel studies, the evidence of long-run PPP tends to be much more favorable when high inflation countries are included. Moreover, the interpretation of the panel studies' results is not always very obvious. For example, Karlsson and Löthgren (2000) using Monte Carlo simulations, found that for panels with long spans of data, the null hypothesis of unit roots can be erroneously rejected even when only a small proportion of the series is stationary. For panels with short spans of data, however, Karlsson and Löthgren (2000) found that the null hypothesis of unit roots 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 in panel unit root tests 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 inclusive results.

Still other researchers, using different empirical methodologies, found mixed results with data from the post-Bretton Woods period. For example, Cheung and Lai (1998), 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 (1992) where stationarity is the null, rather than the alternative hypothesis, found favorable evidence to support long-run PPP. However, a recent study by Caner and Kilian (1999) shown that severe size distortion can result with the use of conventional asymptotic critical values for tests of the null hypothesis of stationarity if the model under the null hypothesis is highly persistent. On the other hand, using size-adjusted critical values can overcome the problem of size distortions, but result in low power of the tests for economically plausible values of the first-order autoregressive (i.e., AR (1)) parameter. Finally, Baum, Barkoulas, and Caglayan (1999), allowing for fractional differencing or structural breaks, found no evidence to support long-run PPP, however.

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

3. Testing For Unit Root

In this paper, we use two of statistical methods to study long-run PPP using data from the post-Bretton Woods era for twenty-one real exchange rates. We start with a more efficient univariate unit root test proposed by Elliott, Rothenberg, and Stock (1996), and Elliott (1998). Next, we allow for a more flexible model structure, such as shifts in the mean or the time trend, for those real exchange rate series that we cannot reject the null hypothesis of a unit root using the more efficient unit-root tests. Since we are using only univariate time series of the real exchange rates, we avoid the potential problems associated with using panel data. Furthermore, since our data are from the post-Bretton Woods period, we also avoid the criticisms of using long spans of data that mixed both fixed and floating exchange rates data.

Previous studies on long-run PPP using efficient unit-root test include Cheung and Lai (1998), Caner and Kilian (1999), and Kuo and Mikkola (1999). Cheung and Lai (1998), using data from the post-Bretton Woods period, examined ten real exchange rates using the efficient unit-root test of Elliot et al (1996). Kuo and Mikkola (1999) studied only the U.S./U.K. real exchange rate using data spanning 134 years, thus mixing both fixed and flexible exchange rate periods. The main purpose of the study by Caner and Kilian (1999) was to study the size and power of tests that have stationarity as the null hypothesis. They use the efficient unit-root test of Elliot et al (1996) to study long-run PPP as a comparison to those tests. Our paper is a much more thorough study of long-run PPP using a larger data set, and using the efficient unit-root tests of both Elliot et al (1996) and Elliot (1998), in addition to using tests that allow for a more flexible model structure.

To be specific, we define 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 Elliot et al (1996) and Elliot (1998) are similar, differing only in the initial condition assumption. We will review briefly the tests and point out the differences. The efficient unit-root test of Elliot 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, Elliot et al (1996) derived the asymptotic maximal power envelope. From the power calculations, Elliot 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 $H_A : \delta_0 < 1$.

Elliot's (1998) efficient unit-root test, denoted as DF-GLS^u, differs from Elliot et al (1996) in its assumption about the initial value of the alternative model. Specifically, both Elliot et al (1996) and Elliot (1998) assume that their data $(y_1, y_2, \ldots, y_T)$ 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. Elliot et al. assumed that the initial value of u_t, i.e., u_0 is zero both when $\rho = 1$ and when $\rho < 1$, so that $u_1 = v_1$. Elliot (1998) 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 $\frac{Var(v_t)}{(1-\rho^2)}$. Since the alternate assumption involves the unknown parameter ρ, Elliot (1998) 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 Elliot et al (1996), and a different set of the critical values of the test statistics are derived in Elliot (1998). To implement the DF-DLS[u] test, equation (3) is estimated by least squares, with q_t^d, z_t, and β are as defined before, except that now $\tilde{q}_t = [(1-\bar{\rho}^2)^{1/2} q_1, (1-\bar{\rho}L)q_2, \dots, (1-\bar{\rho}L)q_T]'$, and $\tilde{z}_t = [(1-\bar{\rho}^2)^{1/2} z_1, (1-\bar{\rho}L)z_2, \dots, (1-\bar{\rho}L)z_T]'$.

In practice, however, it is difficult to know whether the sample data conform to the data generating models of Elliott et al (1996) or Elliott (1998). We report test results for both in our paper. Following Elliot 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 Elliot et al (1996). For the DF-GLS[u] test, we also follow Elliot (1998) and use $\bar{c} = -10$ in both test with a constant, and test with a constant and a linear trend. We discuss our data set and present our empirical results in the next section.

4. Empirical Results

The source of our data is the OECD G-7 countries, supplied on a diskette. Our data consist of monthly observations from April 1973 to February 1999 for the G-7 countries, 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 only bilateral nominal exchange rate available on the diskette uses the U.S. dollar as the base currency, i.e., foreign currency per U.S. dollar. Since, we are also interested in whether the use of non-U.S. dollar based real exchange rates may produce different results, as other studies have found, we therefore 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.[3]

We start our empirical tests by first presenting in Table 1 our test for unit-root using the ADF test. This is motivated by two factors. First, our results 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. Second, the ADF unit-root test results will provide a comparison to the efficient unit-root tests of Elliott et al (1996), and Elliott (1998). 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, call this ***lmax***. We test the statistical significance of this pre-determined maximum lag using the conventional t-statistic at the 10% significance level. If it is not statistically significantly different from zero, this lag is dropped and we re-estimate Equation (7) with ***l = lmax-1***. This process is repeated until the last included lag is statistically significant at the 10% significance level. We start with an upper bound for ***lmax*** = 12, and a lower bound of ***lmax*** = 1. If, however, the lag length determined is the same as the upper bound ***lmax*** value, we start over with ***lmax*** = 14. Note that we have included a linear time trend to allow for the possibility that the real exchange rate may be trend-stationary.[4] 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 not 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 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. 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 in results 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 tests. 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 rate with five out of six real exchange rates each. For the lira-based real exchange rates, we have only four out of six cases where the real exchange rates are either stationary or trend-stationary at the 10% significance level or better. Overall, we conclude that there are no major differences in results between U.S. dollar-based or non-U.S. dollar-based real exchange rates, with perhaps the Italian lira-based real exchange rates as the exceptions.

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. 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. Since we are looking for evidence of long-run PPP in the post-Bretton Woods era, we therefore accept the more favorable results provided by the DF-GLS test. This leaves three real exchange rates that long-run PPP does not appear to hold – the Canadian dollar/U.S. dollar, the sterling/lira, and the mark/lira rates. We investigate these three real exchange rates further in the next section.

5. Structural Breaks and Long-Run PPP

In this section, we use a more flexible ADF regression by allowing for the possibility of structural breaks in the real exchange rate time series. We start with the simplest structural-break model where there is only one break point, and further limit our attention to the class of models where the breakpoint is endogenously determined. The first structural-break model that we consider is the "innovational outlier model" of Perron (1989, 1997):

$$q_t = \mu_0 + \mu_1 DU_t + \gamma_0 t + \gamma_1 D(TB)_t + \phi q_{t-1} + \sum_{i=1}^{l} \lambda_i \Delta q_{t-1} + \xi_t \,, \tag{8}$$

where ***TB*** is the single break-point in the time series, $DU_t = 1$ for ***t > TB***, and 0 elsewhere; $D(TB)_t = 1$ for ***t = TB + 1***, and 0 elsewhere; and ξ_t is a serially uncorrelated error process. The null hypothesis of a unit-root is $H_0 : \phi = 1$. The lag length for Equation (8) is chosen by the same general-to-specific method described earlier. The break-point is determined by using the maximum of the absolute value of the t-statistic on μ_1.[5] A trim value of 15% is also used at the beginning and the end of the data series to rule out the possibility of the break-point occurring at those two end points. Our results for the remaining three real exchange rates are presented in Table 3.

Starting with the sterling/lira real exchange rate, the break-point date identified is 1985:06. We have no ready explanation for this break-point date. The null hypothesis of a unit-root is barely not rejected at the 10% significance level (-4.533 vs. –4.58 at the 10% critical significance level). For the mark/lira exchange rate, the break-point date identified is 1992:07. This seems reasonable since it is sufficiently close to the exchange rate crisis of September 1992 when the lira was devaluated by about 15% against the mark because of speculative attacks on the lira and other European currencies belonging to the Exchange Rate Mechanism. The mean-break term (***DU***) is highly significant and its sign is consistent with a devaluation of the lira against the mark. Moreover, the null hypothesis of a unit root is rejected at the 5% significance level (-4.897 vs. –4.80 at the 5% critical significance level). Finally, the break-point date of 1986:05 for the Canadian dollar/U.S. dollar is also reasonable since it is within the period of the sharp real depreciation of the U.S. dollar between 1985 and 1987. The sign of the mean-break term is also consistent with this explanation, although we cannot assess its statistical significance since we cannot reject the null hypothesis of a unit root at the 10% and the 5% significance levels.

We have tried more complicated structural-break models for the sterling/lira and the Canadian dollar/U.S. dollar real exchange rates. For example, we have allowed for a simultaneous break in the intercept and the time trend (Perron, 1997), two breaks in the intercept (Clemente, Montañés, and Reyes, 1998), two breaks in the intercept and the time trend or combination of the two (Lumsdaine and Papell, 1997), but found no significant improvement over the simpler structural break model given by Equation (8). These results are thus not reported. We are somewhat surprised by these results especially with the models with two structural breaks, since, a prior, we believe that they are good candidates for the Canadian dollar/U.S. dollar real exchange rate given the behavior of the U.S. dollar real exchange rate in the 1980s discussed earlier.

In sum, using a more flexible ADF specification by allowing for structural breaks turned up one additional case of long run PPP for the mark/lira real exchange rate. A case could also be made for the sterling/lira real exchange rate where the null hypothesis of a unit root is barely not rejected at the 10% significance level. We are, however, unable to uncover any evidence in favor of long-run PPP for the Canadian dollar/U.S. dollar real exchange rate.

6. 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 (1998) 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 Elliot et al (1996) DS-GLS test. Using a flexible ADF regression by allowing for the possibility of structural breaks, we are able to find evidence of long-run PPP for one additional real exchange rate. In all, we find supportive evidence for long run PPP for nineteen out of twenty-one real exchange rates.

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 using 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.

Table 3
Structural Break Models

Coefficient of:	Exchange Rate Sterling/Lira	 Mark/Lira	 Canadian Dollar/U.S. Dollar
Constant	-0.666 (4.521)	-0.524 (4.901)	0.002 (1.146)
DU	0.026 (3.897)	-0.027 (4.813)	-0.009 (2.967)
t	-1.294E-004 (3.668)	1.743E-004 (4.795)	7.579E-005 (3.279)
D(TB)	-0.061 (2.856)	0.022 (1.162)	0.023 2.078)
q_{t-1}	0.913 (47.708)	0.924 (60.140)	0.966 (91.950)
Lag length:	12	2	13
t-statistic for $H_0 : \phi = 1$	-4.533	-4.897*	-3.238
Break date:	1985:06	1992:07	1986:05

Notes: * Denotes the rejection of the null hypothesis at the 5% significance level.
The absolute value of the t-statistic is given in parenthesis below the estimates.

Footnotes

[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).

[3] 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.

[4] 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, some researchers, e.g., Papell and Theodoridis (1998), consider a linear time trend in the real exchange rate as inconsistent with long-run PPP.

[5] Alternative methods for determining the break-point as suggested in Perron (1997) produce similar results.