WORKING PAPER 30

The Great Appreciation, the Great Depreciation, and the Purchasing Power Parity Hypothesis

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Editorial

On April 3 -4, 1998 the Oesterreichische Nationalbank hosted a joint Euroconference with the CEPR on „Real Exchange Rates: Recent Theories and Evidence". A number of papers presented at this conference is being made available to a broader audience in the Working Paper series of the Bank. This volume contains the fourth of these papers. The first three papers were issued as Working Papers 27-29.

ABSTRACT

Although there has been much recent work on Purchasing Power Parity (PPP), neither univariate nor panel methods have produced strong rejections of unit roots in U.S. dollar real exchange rates for industrialized countries during the post-1973 period. We investigate the hypothesis that these non-rejections can be explained by one episode, the large appreciation and depreciation of the dollar in the 1980s, by developing unit root tests which account for this event and maintain long-run PPP. Using panel methods, we can reject the unit root null for those countries which adhere to the typical pattern of the dollar's rise and fall.

July 3, 1998

Note: The views expressed in this Working Paper are strictly those of the authors and do not, in any way, commit the Oesterreichische Nationalbank nor the CEPR.

I. Introduction[1]

Purchasing Power Parity (PPP) is one of the most enduring topics in international economics, and the question of whether PPP holds during the post-Bretton-Woods system of flexible nominal exchange rates has been extensively analyzed. While the failure of PPP to hold in the short run was obvious after the first few years of generalized floating, long-run PPP has been subject to a "mean reversion in economic thought" (Lothian and Taylor, 1997). In the mid-1970s, models such as Dornbusch (1976) routinely used PPP as a long-run equilibrium condition. By the mid-1980s, the widespread failure to reject unit roots in real exchange rates led authors such as Stockman (1990) to construct models where long-run PPP did not hold. By the mid-1990s, however, research on both long-horizon data and on panels of post-1973 real exchange rates has led to a renewed belief in the validity of long-run PPP.

All variants of PPP postulate that the real exchange rate reverts to a constant mean. Evidence of long run PPP can be provided by tests of a unit root in the real exchange rate. If the unit root null hypothesis can be rejected in favor of a level stationary alternative, then there is long-run mean reversion and, hence, long-run PPP. While a "weak form" of PPP is often investigated by testing for cointegration between the exchange rate, domestic price level, and foreign price level, PPP also requires symmetry between domestic and foreign prices and proportionality between relative prices and the exchange rate.[2]

The starting point for research on PPP during the current float is the observation that, using conventional Augmented-Dickey-Fuller (ADF) tests on univariate real exchange rates for industrial countries, the unit root null is rarely rejected. While these findings were initially taken as evidence against PPP, it has become clear that they say more about the power of unit root tests than about PPP.

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Froot and Rogoff (1995) show that if the real exchange rate follows a stationary AR(1) process and the half life of a PPP deviation is three years, it would take 72 years of data to reject the unit root using a 5 percent Dickey-Fuller critical value. Lothian and Taylor (1997) provide more extensive power simulations with sample moments chosen to match actual long-horizon real exchange rates and conclude that, for the sterling/dollar rate, even a century of data is not sufficient to reject the unit root hypothesis.

In response to these problems, research on long-run PPP has progressed in two directions. First, univariate techniques have been applied to long-horizon real exchange rates spanning one to two centuries. This data, however, necessarily combines periods of fixed and floating nominal exchange regimes, and cannot answer the question of whether evidence of PPP would be found with the same time span of flexible rates. Second, tests for unit roots in panel data, notably those of Levin, Lin, and Chu (1997) and Im, Pesaran, and Shin (1997), have been used to test for PPP among industrialized countries in the post-1973 period.[3]

Panel unit root tests have not produced strong evidence of PPP for U.S. dollar based real exchange rates. Papell (1997), using quarterly data for 21 industrialized countries from 1973 to 1994, cannot reject the unit root null at the 10 percent level when serial correlation is taken into account in calculating lag lengths and computing critical values. This result is unchanged when the sample is extended through 1996 in Papell and Theodoridis (1997a). O'Connell (1998), emphasizing contemporaneous correlation, also reports non-rejections of the unit root null for post-1973 real exchange rates.[4]

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[2] Dornbusch (1987), Froot and Rogoff (1995), and Rogoff (1996) survey the various concepts of PPP.
[3] Another direction for research has been to use unit root tests with more power, notable the DF-GLS test of Elliot, Rothenberg, and Stock (1996). Application of these tests to post-1973 real exchange rates by Cheung and Lai (1997), however, produces only weak additional rejections of the unit root null among industrialized countries.
[4] Papell (1997) discusses how early claims of strong rejections of unit roots in real exchange rates with panel methods were caused by inappropriately imposing homogeneous intercepts, neglecting serial correlation in calculating critical values, inclusion of a time trend, and/or fortuitous selection of countries to include in the panel.

These non-rejections are not caused by the low power of panel unit root tests. Levin, Lin, and Chu (1997) and Bowman (1997) both report very high size adjusted power for panels of the size, time span, and half-lives of the post-1973 real exchange rates. In addition, the same tests have produced

strong evidence of PPP for German mark based real exchange rates. Jorion and Sweeney (1996) rejects the unit root null at the 1 percent level for a panel of European countries, and Papell (1997) reports equally strong rejections for panels of both European and industrialized countries.

Why are the rejections of the unit root hypothesis so much stronger when the German mark, rather than the U.S. dollar, is used as the numeraire currency? Jorion and Sweeney (1996) suggest two explanations: dollar-based real exchange rates are more volatile than mark-based real exchange rates and the European countries which comprise most of the sample are closer to Germany than to the United States. While Papell and Theodoridis (1997b) provide evidence that distance and volatility can help explain the *relative* strength of the rejections of unit roots among panels with different numeraire currencies, these explanations cannot account for the non-rejections of the unit root null with the dollar as numeraire.

Lothian (1997) has proposed a potentially more compelling explanation. The behavior of the dollar, but not the mark, has been dominated by one episode: the large appreciation and depreciation of the dollar in the 1980s. This is illustrated in Figure 1, which shows the real mark/dollar exchange rate and an (unweighted) average of the real exchange rates of the dollar against currencies of 20 industrialized countries.[5] Both of these exchange rates are clearly dominated by the single episode of appreciation and depreciation. Figure 2 depicts the same information for the mark: the real dollar/mark rate and the average real exchange rate for the mark. This picture is very different. The variation in the average real mark exchange rate comes from a number of episodes across the sample and is not dominated by the single episode.

[5] The precise definition of the real exchange rate is provided in Section II, and the countries are listed in Table 1.

This is an example of a more general problem. The results in Levin, Lin, and Chu (1997), Im, Pesaran, and Shin (1997), and Bowman (1997) depend crucially on the assumption of independence across individuals, and are not applicable if cross-sectional correlation is present. For the dollar-based real exchange rates, which are highly contemporaneously correlated, including more countries provides less additional information than for the mark-based rates, lowering the power of the tests. O'Connell (1998) stresses the importance of contemporaneous correlation for panels of real exchange rates, and shows that it cannot be removed simply by subtracting cross sectional means from the data. He proposes a maximum likelihood procedure to account for contemporaneous correlation, but at the cost of severely restricting the degree of serial correlation.

We propose univariate and panel unit root tests, which account for the appreciation and depreciation of the dollar. The motivation for these tests comes from research on testing for unit roots in the presence of structural change. In Perron (1989), the Great Crash of 1929 and oil price shock of 1973 were taken as exogenous events which were not realizations of the underlying data generating mechanism, and changes in the intercept, slope, or both were modeled by dummy variables. In this paper, we make a similar assumption for the great appreciation and depreciation of the dollar in the 1980s. We base this assumption on both economic and statistical reasons. First, despite the huge amount of research on models of exchange rate determination, there has been no successful explanation of the magnitude and duration of the dollar's appreciation in 1980-84 based on economic fundamentals.[6] Second, the magnitude of the rise and fall of the dollar from 1980 to 1987 dwarfs the other post-1973 experiences.

[6] Frankel and Froot (1990), for example, suggest that the dollar "overshot the overshooting equilibrium."

The behavior of dollar-based real exchange rates under the current float is not well suited for modeling by existent tests for unit roots in the presence of structural change. As shown in Figure 1, the dollar depreciated from 1973-79, appreciated sharply during 1980-84, depreciated even more sharply in 1985-87, and has fluctuated thereafter. Tests that allow a one-time break in the intercept, slope, or both do not capture these swings. In addition, these tests do not impose the purchasing power parity requirement of a constant mean in the long run.

The specific test which we propose allows for three changes in the slope, but no breaks in the intercept, of the trend function for the real exchange rates. This extends Perron's "changing growth" model in four ways: First and most obviously, we allow for three breaks instead of one. Second, the breaks are determined endogenously. Searching for three breaks with approximately 100 observations raises computational questions, especially for the calculation of critical values, that we discuss below. Third, the coefficients on the time trend and the dummy variables which depict the breaks are constrained to produce a constant mean following the final break, making the restrictions consistent with PPP. Fourth, the tests are developed in a panel, as well as a univariate, context.

The intuition that motivates the tests is to treat the rise and fall of the dollar in the 1980s as being determined outside the data generating process. In theory, the first break would pick up the start of the dollar's appreciation at the end of the 1970s, the second break would be caused by the switch from appreciation to depreciation in 1984-85, and the third break would depict the end of the depreciation in 1987. In practice, it is important to understand that, because of the endogenous determination of the break dates, there is absolutely nothing to constrain the breaks to be anywhere near the dates mentioned above.

We first use univariate methods, subject to the restrictions described above, to test the unit root null against the "broken trend" alternative for the real exchange rates of 20 industrialized countries with the dollar as the numeraire currency. Given the lack of power in univariate unit root tests, we are not surprised that these tests provide little evidence of PPP. The unit root null can be rejected (at the 5 percent level) for only one country, about the same as with ADF tests that do not incorporate breaks.

We then proceed to develop panel unit root tests in the presence of "PPP restricted" structural change. Since the dates of the breaks are constrained to be the same across the different countries' real exchange rates, we choose a subset of the 20 countries for which the assumption of common break dates seems tenable. With the univariate methods, 15 of the 20 real exchange rates exhibit the "typical" pattern of the first break (appreciation of the dollar) around the end of the 1970s, the second break (depreciation) in late 1984 or early 1985, and the third break (end of the depreciation) around 1987.

Using panel methods, we reject the unit root null hypothesis in favor of the PPP restricted broken trend alternative at the 5 percent level, providing strong evidence that unit roots in post-1973 real exchange rates can be rejected once the great appreciation and depreciation of the dollar in the 1980s is taken into account. Three more tests provide additional evidence that the rise and fall of the dollar is the cause of the non-rejection of the unit root hypothesis in conventional panel tests. First, using conventional panel unit root tests with an alternative of level stationarity, which do not incorporate structural change, we cannot reject the null for the panel of 15 countries. Second, constructing panels of 16 – 20 countries by adding "atypical" countries to the panel of 15, the strength of the unit root rejections falls much faster as the size of the panels increases when structural change is incorporated. Third, constructing panels of 11 – 14 countries by subtracting "typical" countries from the panel of 15, the strength of the unit root rejections falls much faster as the size of the panels decreases when structural change is not incorporated.

The paper is organized as follows: Univariate tests for unit roots in the presence of restricted structural change are presented in Section II. These tests do not provide much evidence of purchasing power parity. Panel tests are developed in Section III. Application of these tests to panels with similar break dates provides much stronger evidence of PPP. Conclusions are presented in Section IV.

II. Univariate tests

The purpose of this paper is to investigate the purchasing power parity hypothesis among industrialized countries in the post-Bretton-Woods flexible exchange rate period. We use quarterly, nominal, end-of-period exchange rates and Consumer Price Indexes for industrialized countries, obtained from the International Monetary Fund's International Financial Statistics (CD-ROM for 3-97). The data start in the first quarter of 1973 and end in the fourth quarter of 1996, providing 96 quarterly observations. There are 23 countries that are considered industrialized by the IMF. We do not use data for Iceland because of the existence of gaps in its CPI and for Luxembourg because it has a currency union with Belgium. The 21 remaining countries provide 20 real exchange rates with the U.S. dollar as the numeraire currency.

The real (dollar) exchange rate is calculated as follows,

$$q = e + p^* - p \ , \qquad (1)$$

where q is the logarithm of the real exchange rate, e is the logarithm of the nominal (dollar) exchange rate, p is the logarithm of the domestic CPI, and p^* is the logarithm of the U.S. CPI.

The most common test for PPP is the univariate ADF test, which regresses the first difference of a variable (in this case the logarithm of the real exchange rate) on a constant, its lagged level and k lagged first differences,

$$\Delta q_t = m + a q_{t-1} + \sum_{i=1}^{k} c_i \Delta q_{t-i} + e_t \ , \qquad (2)$$

A time trend is not included in equation (2) because such an inclusion would be theoretically inconsistent with long-run PPP. The null hypothesis of a unit root is rejected in favor of the alternative of level stationarity if α is significantly different from zero. We use the recursive t-statistic procedure proposed by Hall (1994) to select the value of k, with the maximum value of k equal to 8 and the ten percent value of the asymptotic normal distribution used to determine significance.[7] The unit root null can be rejected (at the 5 percent level) for only one (the United Kingdom) out of 20 real exchange rates.[8]

A general principle of unit root tests, emphasized by Campbell and Perron (1991), is that nonrejection of the unit root hypothesis may be due to misspecification of the deterministic trend. Perron (1989) formalized this by allowing for a one-time exogenously determined break in the intercept, time trend, or both of the trend function, and the methodology has been extended to allow the breaks to be determined endogenously. These methods, however, are not suitable for testing purchasing power parity because the alternative hypothesis, (broken) trend stationarity, is not consistent with PPP.

Tests for a unit root in non-trending data which allow one break in the intercept, as in Perron and Vogelsang (1992), are consistent with a weaker version of PPP, called “qualified” PPP by Dornbusch and Vogelsang (1991) and “quasi” PPP by Hegwood and Papell (1998). The two-break unit root tests of Lumsdaine and Papell (1997) could also be extended to non-trending data. Post-1973 real exchange rates, however, do not appear to be well modeled by such tests. As shown in Figure 1, these exchange rates are not obviously characterized by a one (or two) time change in the intercept.

We extend Perron's “changing growth” model, which allows a one-time change in the slope, but not in the intercept, of the deterministic trend in a way that is consistent with PPP. As described by Perron (1989,1997), since the changes in slope are presumed to occur rapidly, the model is of the additive outlier type and is estimated by a two-step procedure. First, the series is detrended using the following regression,

$$q_t = \mu + \beta t + \gamma_1 DT1_t + \gamma_2 DT2_t + \gamma_3 DT3_t + z_t \,. \qquad (3)$$

The unit root null is rejected if α is significantly different from zero in the regression,

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[7] As discussed by Campbell and Perron (1991) and Ng and Perron (1995), this procedure has better size and power properties than alternative methods such as selecting k based on the AIC or BIC.

[8] This result is so well known that we do not report the details.

$$\Delta z_t = a z_{t-1} + \sum_{i=1}^{k} c_i \Delta z_{t-i} + e_t \ . \tag{4}$$

The breaks occur at times TB1, TB2, and TB3, and the dummy variables DTI_t = (t - TBI) if t > TBI , 0 otherwise, I = 1,... , 3. This unit root test becomes a test of purchasing power parity by the addition of the restriction,

$$\beta + \gamma_1 + \gamma_2 + \gamma_3 = 0, \tag{5}$$

which imposes a constant mean following the third break.

There are two possible methods for endogenously the selecting the break dates: by choosing the breaks which minimize the t-statistic on α in (4) and by choosing the breaks which maximize the F-statistic on DT1, DT2, and DT3 in (3). We use the second method for two reasons. First, since we want to investigate the proposition that the nonrejection of unit roots in real exchange rates is caused by the rise and fall of the dollar, we want the trend to fit the data as closely as possible. Second, maximizing the F-statistic in (3) is computationally less burdensome, which becomes important for calculating the critical values.

Choosing three breaks endogenously forces some compromises in calculating critical values. It was not possible, using available computers, to directly extend the Monte Carlo methods described above to the panel case.[9] Instead, we first choose the break that

produced the best fit for the following equation,

$$q_t = \mu + \beta t + \gamma_1 DT1_t + z_t \ , \tag{6}$$

[9] The model with 96 observations takes about 20 minutes to estimate on a Pentium Pro 200, which would require 10 weeks to compute 5000 replications. These calculations, in contrast, took less than one day. We computed 1000 replications for the model where the three breaks were chosen globally, and the critical values were very close to those that we report.

without imposing the PPP restriction. Fixing the first break, we choose the second break, again without imposing the PPP restriction that produced the best fit. Then, fixing the first two breaks, we choose the third break, this time with the PPP restriction, that produced the best fit. Finally, again with the PPP restriction, we set TB1 = TB2 and TB2 = TB3, chose the "new" TB3 with the best fit, and repeat the process 10 times. Once the break dates are chosen, the series are detrended and the unit root test statistics calculated as in Equations (3) and (4).

We calculate critical values using Monte Carlo methods with randomly generated data. First we fit autoregressive (AR) models to the first differences of the data, using the Schwarz criterion to choose the optimal AR model. Then we use the optimal AR model in order to generate the errors for our data. We use the optimal AR model with iid $N(0,\sigma^2)$ innovations to construct pseudo samples of size equal to the actual size of our series (96 observations). The critical values for the finite sample distributions are taken from the sorted vector of 5000 replicated statistics. The critical values are about 15 percent higher, in absolute value, than the comparable critical values in Perron (1997) because we allow for three breaks (subject to the PPP restriction), rather than one, in the trend function, and are more than 50 percent greater than critical values of ADF tests for trending data without breaks.[10]

The results of the univariate unit root tests, reported in Table 1, provide little evidence of PPP. While the unit root null can be rejected at the 5 percent level for Switzerland, it cannot be rejected, at even the 10 percent level, for any other country. This should not be surprising. Since univariate ADF tests have very low power in samples of this size and time span, there is no reason to believe that univariate tests for unit roots in the presence of structural change would have substantially

[10] While we did not want to calculate 21 sets of critical values (one for each series), we did some experimentation with computing critical values based on AR models of the first differences of the actual data. As in Papell (1997), this made little difference for post-1973 real exchange rates.

greater power. We also performed the univariate test for the (unweighted) average of all 20 real exchange rates, and found the same non-rejection of the unit root null.[11]

The dates of the breaks and the coefficients on the dummy variables are reported in Table 2. The "typical" pattern of breaks, reflecting the rise and fall of the dollar in the 1980s, is shared by the real exchange rates of 15 of the 20 countries, those of Austria, Belgium, Denmark, Finland, France, Germany, Ireland, Italy, Japan, Netherlands, Norway, Spain, Sweden, Switzerland, and the United Kingdom. The first break, with a positive value for γ_1 indicating appreciation of the dollar, occurs in 1978(IV) - 1980(IV). The second break, with negative γ_1 reflecting depreciation of the dollar, is in 1984(III) - 1985 (II), and the third break, with positive γ_1 signifying the end of the dollar's depreciation, is in 1986(IV) - 1988(I). The coefficient on the time trend β is negative for all 15 countries, reflecting the depreciation of the dollar from the start of generalized floating in 1973 to the end of the 1970s. These findings are illustrated in Figure 3 for a selection of 4 of the 15 "typical" real dollar exchange rates, those for the German mark, Japanese yen, Swiss franc, and British pound.

Canada is the most dramatic exception to the typical pattern, with break dates far removed from those described above and opposite signs on all of the coefficients. It is the only country for which the coefficients on the break dummy variables do not fit the pattern. The third breaks for Greece and Portugal occur in 1991(IV), four years after the others. The real exchange rates for Australia and New Zealand depreciated against the dollar in the 1970s, and their first breaks did not occur until 1981. These findings are depicted in Figure 4 for the Australian dollar, Canadian dollar, Greek drachma, and Portuguese escudo real dollar exchange rates.

In order to provide a more formal measure of how well the individual real exchange rates fit the "typical" pattern, we calculate the root median squared error (RMSE) for each country and report the

[11] Since the chosen value of k was equal to the maximum (8) in a number of cases, we estimated the models with the maximum raised to 12. The results were unchanged.

results in Table 2. The RMSE is defined as the square root of the sum (over the three breaks) of the squared deviations (measured in quarters) from the median values of the breaks[12].

The results of this calculation confirm the choice of typical and atypical countries, with the 15 smallest RMSE's associated with the 15 typical countries. They also confirm the choice of the size of the groups. The largest RMSE in the typical group is for Japan, while the smallest RMSE in the atypical group is for New Zealand. The difference between the RMSE's of New Zealand and Japan is more than 50 percent greater than the difference for Australia (the next smallest of the atypical group) and New Zealand, and more than 300 percent greater than the difference between Spain (the next largest of the typical group) and Japan.

III. Panel tests

The low power of unit root tests against highly persistent alternatives with anything less than a century of data has inspired the development of panel unit root tests which exploit cross section, as well as time series, variation. Variants of these tests have been developed by Levin, Lin, and Chu (1997) (LLC), Im, Pesaran, and Shin (1997) (IPS), Maddala and Wu (1997), and Bowman (1997). Applications of these tests to post-1973 real exchange rates of industrialized countries include Abuaf and Jorion (1990), Frankel and Rose (1996), Jorion and Sweeney (1996), Oh (1996), Y. Wu (1996), O'Connell (1998), Papell (1997), Papell and Theodoridis (1997, 1998), and S. Wu (1997).[13]

[12] The median values of the breaks are 1980(I), 1985(I), and 1987(I). We use the median, rather than the mean, in order to mitigate the effect of outliers and the RMSE, rather than the sum of the absolute value of the deviations, in order to weight one large deviation more heavily than several smaller deviations.

[13] This is an incomplete list, and does not include studies that use data from developing countries, tradable goods prices, panel cointegration, etc.

A panel extension of the univariate ADF test in Equation (2), which accounts for both a heterogeneous intercept and serial correlation, would involve estimating the following equations,

$$\Delta q_{jt} = m_j + a q_{jt-1} + \sum_{i=1}^{k} c_{ij} \Delta q_{jt-i} + e_{jt} \ , \tag{7}$$

where the subscript j indexes the countries, and μ_j denotes the heterogeneous intercept. The test statistic is the t-statistic on α. In Papell (1997), we estimate equation (7) using feasible GLS, with α equated across countries and the values for k taken from the results of univariate ADF tests.[14] For quarterly data with a panel of 20 industrialized countries, we could not reject the unit root null at the 10 percent level with the U.S. dollar as the numeraire currency, but could reject the null at the 1 percent level with the German mark as numeraire.[15]

The non-rejections of unit roots in post-1973 real exchange rates with the dollar as numeraire are not caused by low power of panel unit root tests. In Papell (1997), we estimate a value of α of -0.069 for a panel of 20 industrialized countries. Bowman (1997) reports that, for a panel of 20 members with 100 observations, the 5 percent size adjusted power of the LLC test, with α equaling -0.05 for each member, is 0.99. The opposite results of panel unit root tests with the dollar and mark as numeraire are not caused by serial correlation. While Im, Peseran, and Shin (1997) report a loss of power with serial correlation, and Papell (1997), choosing the value of k by the data dependent methods described above, calculates critical values which are 15 percent larger (in

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[14] Im, Peseran, and Shin (1997) develop tests where ? can vary across countries. Based on the results of univariate ADF tests in Papell (1997), this does not appear to be important for our sample of dollar-based real exchange rates of 20 industrialized countries. Bowman (1997) shows that there is a loss of power of the IPS tests, relative to the LLC tests, when ? is equal across members of the panel.

[15] Papell and Theodoridis (1997b) estimate panels with all 21 industrialized countries as numeraire, and find strong rejections of unit roots in real exchange rates for most of the European, but none of the non-European, countries.

absolute value) than those in Levin, Lin, and Chu (1997), there is no difference in the amount of serial correlation between dollar and mark real exchange rates.[16]

The clearest difference between dollar and mark real exchange rates involves contemporaneous, or cross-sectional, correlation. It is clear from Figures 1 and 2 that dollar real exchange rates are much more highly contemporaneously correlated than mark real exchange rates. In practice, there is a trade-off between accounting for serial and contemporaneous correlation. O'Connell (1998) proposes a maximum likelihood estimator which, given the sample size of quarterly post-1973 data, severely restricts the degree of allowable serial correlation. Papell (1997) uses a feasible GLS estimator, which does not iterate to maximum likelihood because of the presence of a lagged dependent variable, but which allows for more flexibility in modeling serial correlation.[17]

We extend the unit root tests in the presence of restricted structural change, developed above, to the panel context. The dates of the breaks are first chosen by using the following feasible GLS regressions,

$$q_{jt} = \mu_j + \beta t + \gamma_1 DT1_t + \gamma_2 DT2_t + \gamma_3 DT3_t + z_{jt}, \qquad (8)$$

subject to the PPP restriction described in Equation (5), where the dates of the breaks are chosen endogenously to maximize the joint log-likelihood. At this stage, while the intercepts are heterogeneous, the coefficients on the trend and the dummy variables are constrained to be equal across[18] countries.

Once the break dates are chosen, the series are detrended as follows,

$$q_{jt} = \mu_j + \beta_j t + \gamma_{j1} DT1_t + \gamma_{j2} DT2_t + \gamma_{j3} DT3_t + z_{jt}, \qquad (9)$$

[16] Bowman (1997) shows that size adjusted power falls much faster for the LLC test than for the IPS test when only a subset of the members of a panel are stationary. Since rejection of the unit root null is normally interpreted as evidence that all real exchange rates are stationary, we view this as an advantage of the LLC tests.

[17] O'Connell (1998) shows that, if there is no serial correlation or if both the lag lengths and the values of the c's are the same for each country, panel unit root tests of real exchange rates using GLS are invariant to the choice of numeraire currency. These restrictions, however, are rejected by both O'Connell (1997) and Papell and Theodoridis (1997b).

[18] The coefficients on the trend and the dummy variables are equated across countries at this stage to decrease computation time, which becomes necessary for calculating the critical values. We did some experimentation with allowing these coefficients to vary across countries, and the choice of breaks was not affected.

where the coefficients on the trend and the dummy variables are now allowed to vary across countries. The test statistic is the t-statistic on α in,

$$\Delta z_{jt} = a z_{jt-1} + \sum_{i=1}^{k} c_{ij} \Delta z_{jt-i} + e_{jt} \,. \tag{10}$$

As above, the unit root null is rejected against the broken trend stationary alternative if α is significantly different from zero. The dates of the breaks and value of α are constrained to be equal across countries, but the intercepts, coefficients on the trend and dummy variables, and the values of the k's and the c's are heterogeneous. Purchasing power parity under the alternative is imposed by the restriction,

$$\beta_j + \gamma_{j1} + \gamma_{j2} + \gamma_{j3} = 0, \tag{11}$$

for each country. This imposes a constant mean following the third break.

As in the univariate case, choosing three breaks endogenously in the panel context forces compromises in calculating critical values. We first choose the break that produced the best fit for the following set of equations,

$$q_{jt} = \mu_j + \beta t + \gamma_1 DT1_t + z_{jt} \,, \tag{12}$$

without imposing the PPP restriction. Following the procedure for the univariate case, we then choose the second and (imposing the PPP restriction) third breaks, and iterate. Once the break dates are chosen, the series are detrended and the unit root test statistics calculated as in Equations (9) and (10). The critical values for the finite sample distributions are calculated with 5000 replications.[19] Critical values with and without breaks are reported in Table 3. Allowing

[19] The model with 15 countries and 96 observations takes about 80 minutes to estimate on a Pentium Pro 200, which would require 40 weeks to compute 5000 replications. These calculations took between two and three days. We computed 500 replications for the model with 15 countries where the three breaks were chosen globally, and the critical values were very close to those that we report.

for three (PPP restricted) breaks raises the critical values by about 90 percent.[20]

The results of the panel unit root tests in the presence of restricted structural change are described in Table 3, and the break dates and coefficients are reported in Table 4. The unit root null can be rejected at the 5 percent level (with a p-value of .026) for the panel of 15 real exchange rates, which exhibit the "typical" pattern in the univariate tests, providing strong evidence of purchasing power parity. The first break occurs in 1978 (IV), the second in 1985 (I), and the third in 1987 (III). The coefficients on the trend (negative), DT1 (positive), DT2 (negative), and DT3 (positive) are consistent with the general pattern of real depreciation of the dollar in the 1970s, sharp appreciation in the early-to-mid 1980s, and even sharper depreciation until 1987 that was observed in the univariate estimates.[21]

How can we be sure that the non-rejections of unit roots in real exchange rates, reported above, are caused by the rise and fall of the dollar in the 1980s? It is possible that our choice of countries is simply a fortuitous selection, which is particularly favorable to the PPP hypothesis. We investigate this by estimating a "standard" panel unit root model which, as described in Equation (6), does not account for structural change, for the same panel of 15 countries. The results are reported in Tables 3 and 4. The unit root null cannot be rejected at even the 10 percent level, a result which is consistent with the failure to find panel evidence of PPP with the U.S. dollar as the numeraire currency.

[20] A potential problem with the critical values is that they do not incorporate our method of selecting the panel of 15 "typical" real exchange rates out of the group of 20 countries. In order to correct for this, we generated 20 series of 96 observations, calculated three breaks for each series using the procedure described above, chose the 15 series with the smallest RMSE's, and calculated critical values for the panel with those 15 series. These critical values were very close to the critical values for the panel of 15 countries reported in Table 2.

[21] These coefficients are from the estimates in Equation (7), and are constrained to be equal across countries.

A common method for measuring mean reversion in real exchange rates is to calculate the half-lives of PPP deviations, the amount of time that it takes a shock to the real exchange rate to revert 50 percent back to its mean value. The half-life is normally defined as $\ln(.5)/\ln(1+\alpha)$. In the absence of breaks, the half-life for the panel of 15 is 10.65 quarters, or 2.66 years, a number consistent with previous work. With the breaks, the half-life falls to 1.92 quarters, about one-half year. While it is not surprising that removing the largest and most persistent PPP deviation would lower the half-life of real exchange rate deviations, the 82 percent fall in the half-life provides strong evidence of the importance of the great appreciation and depreciation for the failure to find long-run PPP in post-1973 real exchange rates.[22]

Another possibility is that our method of accounting for restricted structural change will provide "evidence" of PPP regardless of whether the real exchange rates of the countries included in the panel experience the typical swings of the 1980s. We estimate panel unit tests, both with and without restricted structural change, for the full panel of 20 countries. As reported in Table 3, the unit root null cannot be rejected, at even the 10 percent level, for either panel. Furthermore, incorporating structural change does not strengthen the evidence against unit roots. The p-values rise from .118 with no breaks to .258 with three restricted breaks, and neither the break dates nor the coefficients are close to those found for either the panel of 15 real exchange rates or for most of the countries with the univariate tests.

We also examine panels of between 15 and 20 real exchange rates. Since Canada is the country with the highest RMSE that most clearly violates the typical pattern, we exclude it and

estimate a panel with 19 countries. The results are similar in some respects to those with the full panel of 20 countries. The unit root null cannot be rejected, with or without accounting for structural change, and the p-values rise once restricted structural change is included in the tests. Only the second

[22] These half-lives provide a complete description of the path of mean reversion only in the absence of serial correlation. With lagged first-differences, mean reversion is not necessarily monotonic and depends on all of the coefficients of the model. Nevertheless, we believe that the relative magnitudes of the half-lives between the models with and without breaks conveys useful information.

break date is within one year of the break dates for the panel of 15 countries. While the coefficient on the trend is nearly the same as with the panel of 20, the coefficients on the dummy variables are similar to those for the panel of 15 countries.

We proceed to construct smaller panels by excluding additional countries in decreasing order of their RMSE's. The results are similar for the panel of 18, which eliminates Portugal, and for the panel of 17, which also excludes Greece. These panels fail to provides evidence of PPP. The unit root null cannot be rejected, with or without accounting for structural change, and the p-values rise once restricted structural change is included. The first and second break dates are now within one year of the break dates for the panel of 15 countries, and the signs of the coefficients are the same as in the panel of 15.

The panel of 16 real exchange rates also excludes Australia. The unit root null can nearly be rejected at the 5 percent level (with a p-value of .054) with three restricted breaks, providing evidence of PPP. In the absence of breaks, the unit root null cannot be rejected at the 10 percent level. All three break dates are within one year of the break dates for the panel of 15 countries, and the signs of the coefficients are again the same as in the panel of 15.

We then examine panels of between 11 and 14 real exchange rates of "typical" countries, again constructed by excluding additional countries in decreasing order of their RMSE's. These panels exclude (in order) Japan, Spain, the United Kingdom, and Switzerland. As the size of the panels falls, the evidence of PPP becomes weaker. The p-values fall monotonically for the models without breaks, and fall (almost) monotonically for the models with three breaks. The evidence of PPP, however, is stronger (with lower p-values) for the models with the three breaks for all four panels.

A useful summary of the results is provided by examining the p-values as the size of the panels is reduced. In the absence of breaks, the p-values generally rise as the number of countries falls, reflecting the lower power of panel unit root tests with smaller panels. With three breaks, the p-values generally fall as the size of the panel decreases from 20 to 15, but then rise as the panels become smaller. Selecting panels of countries for which the breaks are closer to the median provides stronger evidence of purchasing power parity only when the effects of the rise and fall of the dollar are taken into account.

V. Conclusions

The proliferation of recent work on purchasing power parity underscores its importance as a central topic in international economics. The development of panel unit root tests presents both a challenge and an opportunity for researchers attempting to find strong evidence of long-run purchasing power parity using data from the current float. The opportunity occurs because, in contrast with univariate methods, panel unit root tests with 20 individuals and 100 quarterly observations have sufficient power to reject the unit root null in favor of a level stationary alternative, even with half-lives of over four years. The challenge arises because, again in contrast with univariate methods, failure to reject the unit root null in real exchange rates can no longer be ascribed to low power of the tests.

The starting point for this paper is the finding that, once serial and cross-sectional correlation are taken into account, panel unit root tests provide strong rejections of the unit root null when the German mark is used as the numeraire currency, but do not even provide weak rejections with the U.S. dollar as numeraire. While this cannot be explained by differences in serial correlation, differences in cross sectional correlation appear to be more promising. The behavior of the U.S. dollar, but not the German mark, has been dominated by one big event, the rise and fall of the dollar in the 1980s.

We investigate the hypothesis that the failure to reject unit roots in real exchange rates with panel methods can be explained by the great appreciation and depreciation of the dollar. We extend Perron's (1989) "changing growth" model to develop univariate and panel unit root tests that allow for three breaks in the slope of the trend function, with the dates of the breaks determined endogenously. Furthermore, the coefficients on the trend and the break dummy variables are restricted so as to produce a constant mean following the third break. This restriction ensures that rejection of the unit root null in favor of the "broken trend" alternative is evidence of long-run purchasing power parity.

While the univariate tests do not produce evidence against unit roots in real exchange rates, they provide a classification of the 20 countries into two groups. The real exchange rates of 15 of the countries follow the "typical" pattern associated with the rise and fall of the dollar in the 1980s, and these are also the 15 countries with the smallest root median squared error (RMSE) of the break dates. The exceptions to the typical pattern, in ascending order of RMSE's, are New Zealand, Australia, Greece, Portugal, and Canada.

The central result of the paper is that panel unit root tests that account for restricted structural change provide strong evidence against the unit root hypothesis, and thus evidence of purchasing power parity, for the panel of 15 "typical" countries. The unit root null can still be rejected in favor of the restricted structural change alternative if New Zealand is included to form a panel of 16. Unit root tests that do not account for structural change provide no evidence of PPP for these panels. For smaller panels of between 11 and 14 "typical" countries, incorporating structural change increases the evidence of PPP. For larger panels of 17 to 20 countries, the opposite occurs. Incorporating structural change decreases the evidence of PPP.

The premise which motivated this paper was that previous attempts to find evidence of purchasing power parity by applying panel unit root tests to post-1973 dollar-based real exchange rates of industrialized countries have been unsuccessful because the experience has been dominated by one big event: the great appreciation and depreciation of the dollar in the 1980s. By developing panel unit root tests that account for structural change but remain consistent with long-run PPP, we can formally test this hypothesis. Our strong rejection of unit roots in the presence of PPP restricted structural change for the panel of real exchange rates of the 15 "typical" countries characterized by the rise and fall of the dollar, combined with our failure to reject unit roots either in the absence of structural change or for panels which include additional "atypical" countries, provides strong support for this premise.

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Table 1

Univariate Restricted Unit Root Tests

Country	a	**t-statistic**	**p-values**	**k**
Australia	-0.349	-4.47	.384	0
Austria	-0.427	-3.74	.723	4
Belgium	-0.412	-3.96	.631	3
Canada	-0.309	-3.34	.857	6
Denmark	-0.808	-3.98	.620	8
Finland	-0.213	-3.05	.923	7
France	-0.594	-4.70	.284	4
Germany	-0.451	-4.01	.608	4
Greece	-0.521	-4.86	.218	4
Ireland	-0.631	-4.04	.592	8
Italy	-0.302	-3.56	.789	4
Japan	-0.275	-3.75	.720	4
Netherlands	-0.563	-4.51	.367	4
New Zealand	-0.450	-3.94	.641	8
Norway	-0.749	-3.70	.737	8
Portugal	-0.825	-4.63	.315	8
Spain	-0.535	-4.36	.434	8
Sweden	-0.498	-4.33	.451	8
Switzerland	-0.557	-6.02	.015	0
United Kingdom	-0.460	-4.26	.500	8
Average	-0.840	-4.23	.505	8

Critical Values

One Percent	**Five Percent**	**Ten Percent**
-6.18	-5.56	-5.26

Table 2

Univariate Restricted Unit Root Tests
Break Dates and Coefficients

Country	**TB1**	**TB2**	**TB3**	**RMSE**	b	g	g	g
Australia	81(II)	86(II)	88(II)	8.66	.002	.017	-.050	.031
					(2.79)	(8.65)	(-13.40)	(11.92)
Austria	79(III)	85(I)	87(I)	2.00	-.011	.039	-.101	.073
					(-7.86)	(15.09)	(-23.11)	(23.50)
Belgium	79(II)	84(IV)	87(I)	3.16	-.013	.051	-.102	.064
					(-9.31)	(19.32)	(-25.17)	(23.14)
Canada	85(IV)	91(II)	94(II)	37.01	.005	-.016	.033	-.022
					(20.59)	(-21.32)	(21.40)	(-20.31)
Denmark	79(III)	84(IV)	87(II)	2.45	-.010	.042	-.091	.059
					(-8.03)	(17.58)	(-26.22)	(26.03
Finland	79(IV)	84(IV)	87(I)	1.41	-.008	.031	-.069	.046
					(-4.80)	(8.97)	(-12.48)	(12.50)
France	80(I)	84(IV)	87(I)	1.00	-.007	.041	-.092	.058
					(-6.01)	(16.43)	(-23.09)	(22.31)
Germany	79(III)	85(I)	87(I)	2.00	-.006	.036	-.096	.066
					(-4.49)	(14.06)	(-21.86)	(21.13)
Greece	80(I)	84(III)	91(IV)	19.10	-.005	.032	-.045	.017
					(-4.03)	(12.95)	(-20.77)	(21.61)
Ireland	80(III)	85(I)	86(IV)	2.24	-.007	.026	-.076	.057
					(6.00)	(9.55)	(-14.38)	(15.44)
Italy	80(II)	85(I)	86(IV)	1.41	-.002	.024	-0.91	.070
					(-1.83)	(7.41)	(-14.51)	(15.69)
Japan	78(IV)	85(I)	87(I)	5.00	-.017	.032	-0.89	.074
					(-7.80)	(8.79)	(-15.20)	(17.20)
Netherlands	79(II)	85(I)	86(IV)	3.16	-.012	.043	-.103	.072
					(-8.79)	(17.41)	(-22.49)	(21.09)
New Zealand	81(IV)	85(I)	87(III)	7.28	.001	.029	-.079	.049
					(1.02)	(7.46)	(-13.04)	(14.77)
Norway	79(III)	85(I)	87(II)	2.24	-.006	.028	-.066	.044
					(-6.18)	(14.58)	(-22.00)	(21.55)
Portugal	80(III)	83(IV)	91(IV)	19.75	.001	.037	-.058	.021
					(0.43)	(12.17)	(-21.48)	(27.72)
Spain	80(I)	84(III)	88(I)	4.47	-.013	.049	-.081	.045
					(-9.49)	(16.39)	(-22.35)	(23.20)
Sweden	80(II)	84(III)	87(III)	3.00	-.003	.041	-.079	.041
					(-2.52)	(13.11)	(-18.64)	(17.29)
Switzerland	79(I)	85(II)	87(I)	4.12	-.018	.043	-.110	.085
					(-8.60)	(12.32)	(-17.22)	(17.40
United Kingdom	80(IV)	84(IV)	87(IV)	4.36	-.012	.043	-.063	.032
					(-10.18)	(13.74	(-14.67)	(13.86)
Average	79(IV)	85(I)	87(II)		-.007	.031	-.077	.052
					(-7.52)	(17.77)	(-27.34)	(27.45)

Table 3

Panel Unit Root Tests
Three Breaks with PPP Restrictions

Countries	Excluded	a	t-statistic	Critical Values		
				1%	**5%**	**10%**
15	New Zealand	-0.303	-13.80	-14.15	-13.46	-13.06
16	Australia	-0.296	-13.85	-14.66	-13.90	-13.42
17	Greece	-0.255	-13.55	-15.06	-14.30	-13.81
18	Portugal	-0.251	-13.75	-15.57	-14.75	-14.25
19	Canada	-0.210	-13.29	-15.96	-15.19	-14.68
20	None	-0.228	-14.22	-16.45	-15.61	-15.13
14	Japan	-.0294	-12.57	-13.89	-13.07	-12.62
13	Spain	-0.287	-12.04	-13.41	-12.68	-12.20
12	United Kingdom	-0.301	-12.08	-12.87	-12.12	-11.73
11	Switzerland	-0.268	-10.33	-12.47	-11.71	-10.55

No Breaks

Countries	a	t-statistic	Critical Values		
			1%	**5%**	**10%**
15	-0.063	-6.40	-7.81	-7.08	-6.74
16	-0.066	-6.80	-7.94	-7.32	-6.92
17	-0.067	-7.09	-8.18	-7.47	-7.12
18	-0.066	-7.20	-8.37	-7.71	-7.34
19	-0.066	-7.42	-8.66	-7.87	-7.47
20	-0.065	-7.61	-8.81	-8.12	-7.72
14	-0.062	-6.02	-7.67	-6.88	-6.51
13	-0.060	-5.65	-7.38	-6.68	-6.29
12	-0.057	-5.27	-7.09	-6.43	-6.07
11	-0.054	-4.92	-6.88	-6.27	-5.88

Note: The countries that comprise the panel of 20 real exchange rates are listed in Tables 1 and 2. The smaller panels are constı
listed under "excluded".

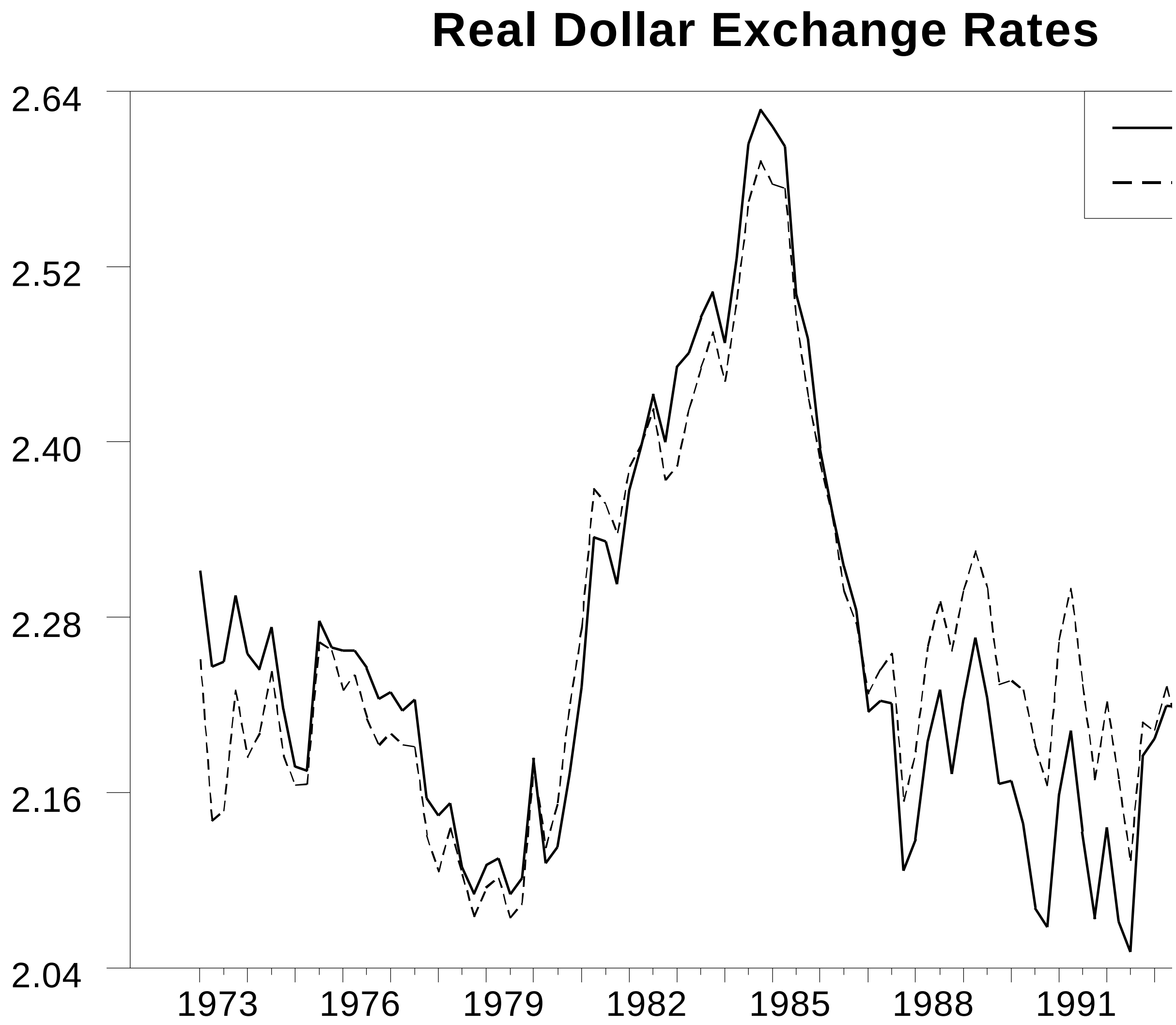
Table 4

Panel Unit Root Tests

Break Dates and Coefficients

Countries	TB1	TB2	TB3	b	g
15	78(IV)	85(I)	87(III)	-.012 (-12.87)	.033 (21.65)
16	80(III)	84(IV)	87(IV)	-.006 (-8.85)	.036 (22.75)
17	80(III)	85(2)	88(IV)	-.004 (-7.07)	.026 (20.25)
18	80(III)	85(2)	88(IV)	-.004 (-6.93)	.026 (20.55)
19	84(I)	85(II)	88(IV)	.005 (13.04)	.030 (8.63)
20	85(II)	92(III)	93(IV)	.006 (26.02)	-.018 (-34.08)
14	80(II)	84(IV)	87(IV)	-.007 (-8.93)	.037 (21.09)
13	80(II)	84(IV)	87(IV)	-.006 (-8.21)	.036 (20.30)
12	78(IV)	85(I)	87(III)	-.011 (-10.02)	.033 (18.83)
11	79(III)	85(I)	87(II)	-.007 (-7.26)	.031 (17.73)

Note: The countries that comprise the panels are listed in Table 3. The t-statistics of the coefficients on the trend and the dumm



Real Dollar Exchange Rates
2.64
2.52
2.40
2.28
2.16
2.04
1973
1976
1979
1982
1985
1988
1991

Figure 1

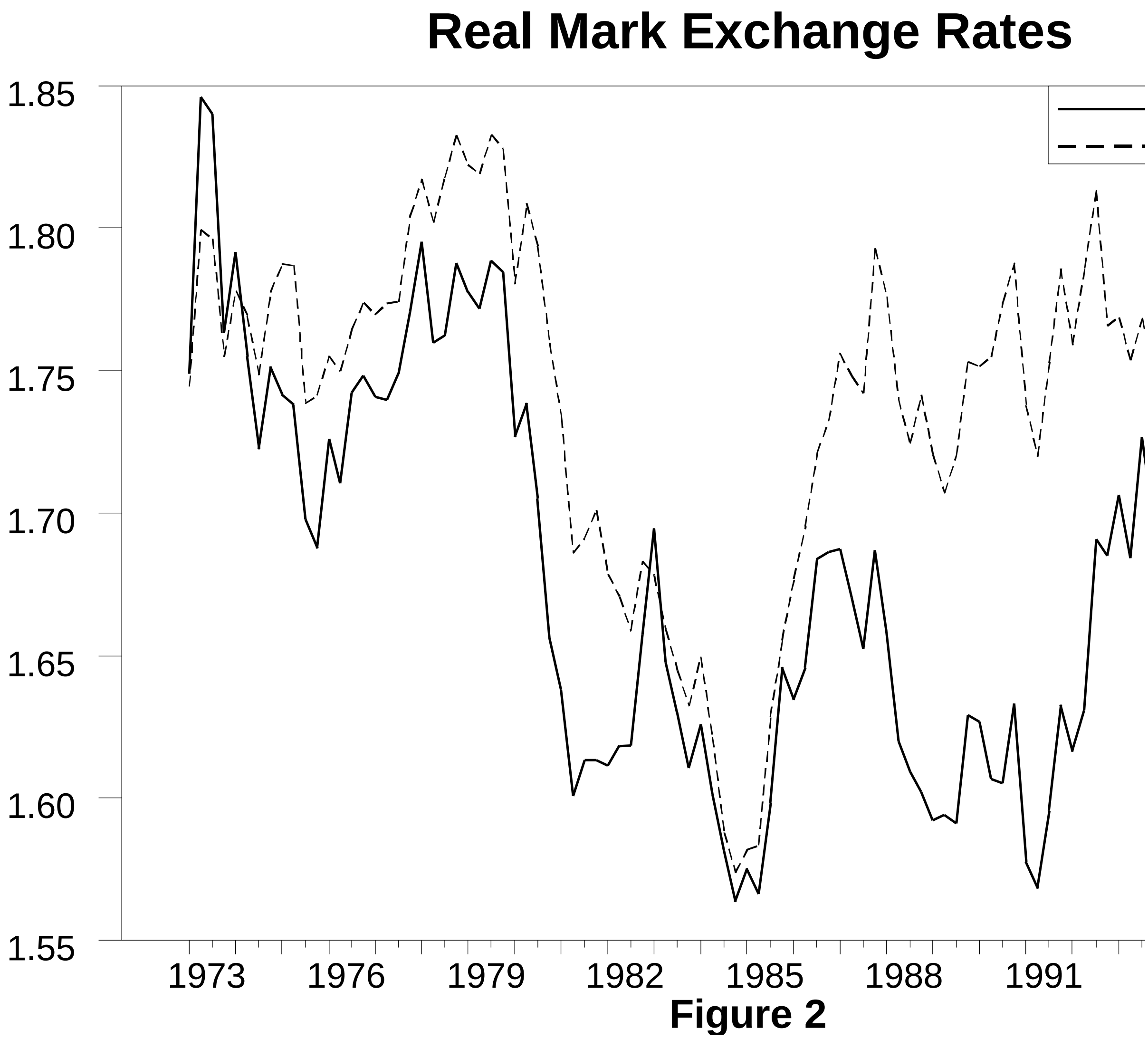

Real Mark Exchange Rates
1.85
1.80
1.75
1.70
1.65
1.60
1.55
1973
1976
1979
1982
1985
1988
1991

Figure 2

Real Dollar Exchange Rates

"Typical" Countries

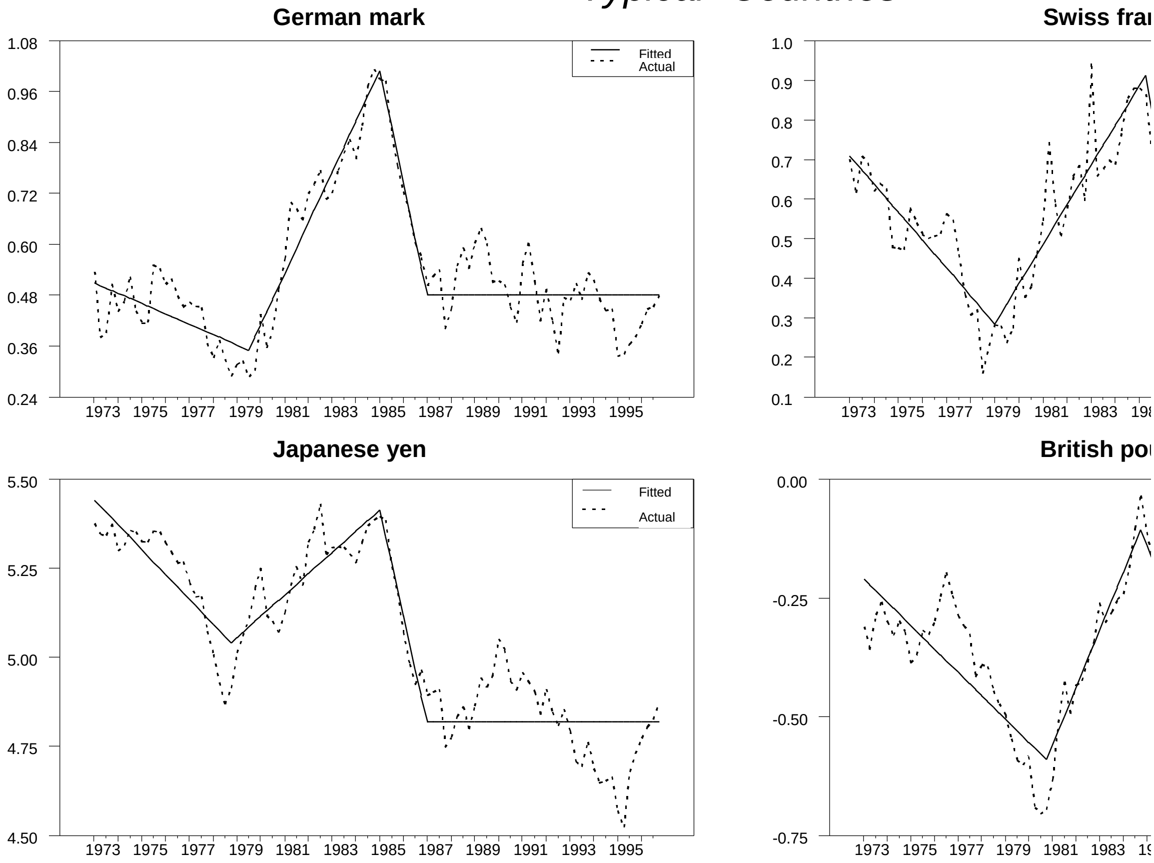


Figure 3

Real Dollar Exchange Rates

"Atypical" Countries

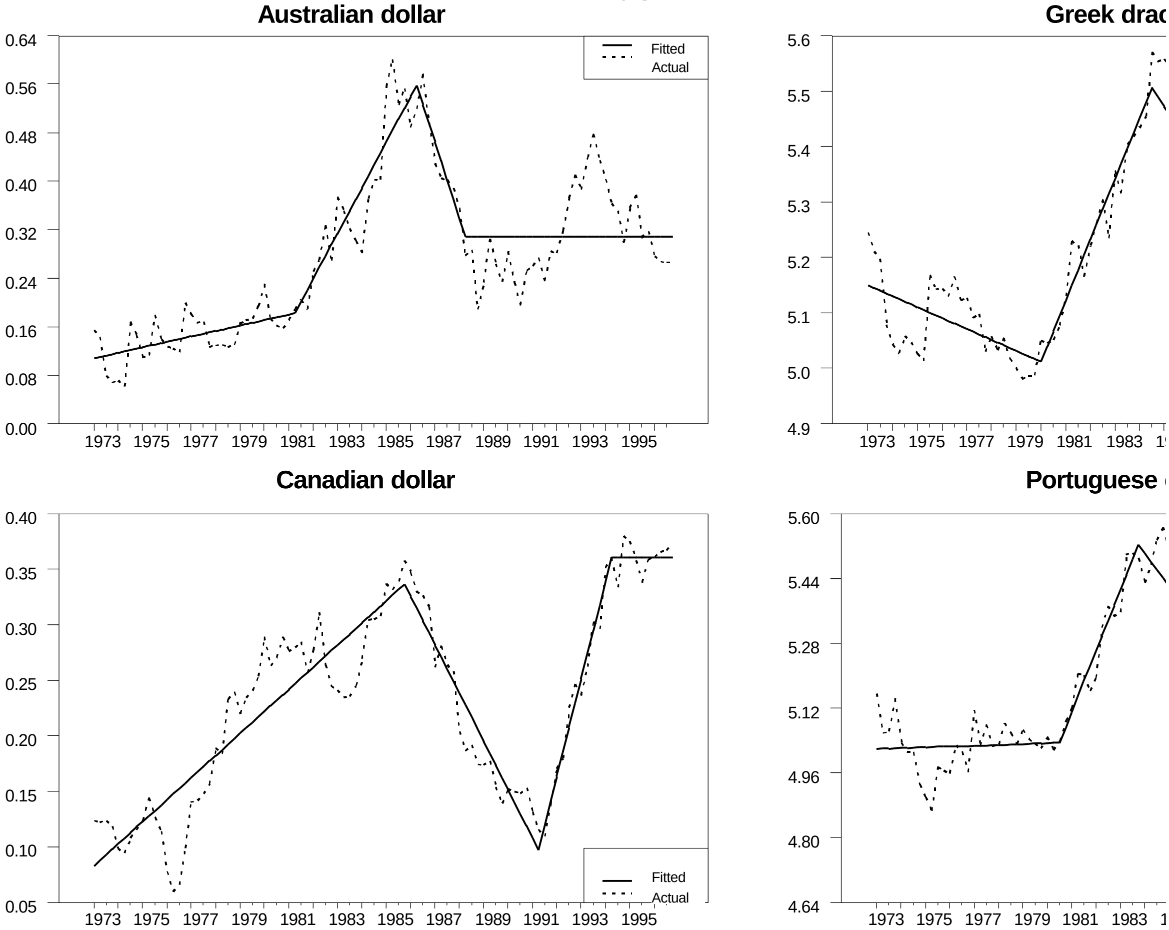


Figure 4