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A panel project on purchasing power parity: Mean reversion within and between countries

Jeffrey A. Frankel[a], Andrew K. Rose[b,*]

[a]*NBER, UC Berkeley, and Institute for International Economics, 11 Dupont Circle, Washington, DC 20036-1207, USA*

[b]*CEPR, NBER, and Haas School, UC Berkeley, Berkeley, CA 94720-1900, USA*

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Abstract

Previous time-series studies have shown evidence of mean-reversion in real exchange rates. Deviations from purchasing power parity (PPP) appear to have half-lives of approximately four years. However, the long samples required for statistical significance are unavailable for most currencies, and are potentially inappropriate because of regime changes. In this study, we re-examine deviations from PPP using a panel of 150 countries and 45 annual post WWII observations. Our panel shows strong evidence of mean-reversion that is similar to that from long time-series. PPP deviations are eroded at a rate of approximately 15% annually, i.e., their half-life is around four years. Such findings can be masked in time-series data, but are relatively easy to find in cross-sections.

Key words: long-run; cross-section; time-series; real; exchange rate; inflation; variation

JEL classification: F30

1. Introduction

Purchasing power parity (PPP) is one of the most important theoretical concepts in international economics. Empirical work on the topic has most

* Corresponding author. Tel. +1 510 642 6609. Fax +1 510 642 4700. E-mail arose@haas.berkeley.edu.

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often used time-series data to compare the percentage changes in bilateral exchange rates with inflation differentials. Many early studies were based on short or medium-length time series, often consisting of post-1973 observations for a few major industrialized countries. They typically did not find strong evidence of PPP.[1] Concerned about inadequate power in their tests, researchers then turned to longer time samples, frequently up to a century or more of time-series data.[2] With longer samples, the evidence has swung back in favor of some long-run tendency toward PPP. At length, consensus has emerged from this literature that there is in fact a moderate tendency for real exchange rates to converge towards a long-run equilibrium. The half-life of PPP deviations appears to be around four years. Froot and Rogoff (1994) provide an excellent critical survey and review of this literature.

This short paper is an empirical re-examination of PPP. Instead of a time-series approach, we use a panel data set of 45 years of post WWII annual data for 150 countries. One motivation of the study is to reiterate the point that the ability to find evidence of PPP depends crucially on the total variation in the data used (including both the number of observations and their variability). A second motivation is to avoid concerns about the use of long time series, since they include potentially serious structural shifts. A typical 100-year or 200-year sample for the pound/dollar rate, for example, includes several shifts between fixed rates, floating, and intermediate regimes. It has been well known since at least Mussa (1986) that real exchange rates behave very differently under different exchange rate regimes. Thus it is reasonable to suppose that the speed of PPP adjustment may also vary with the nature of the exchange rate regimes. Our cross-section approach makes it possible to confine the estimation to the post-1973 period of generalized (dollar) floating, and still have plenty of data for powerful tests.[3]

[1] Examples include: Roll (1979), Frenkel (1981a), Frenkel (1981b),Adler and Lehman (1983), Darby (1981), Mishkin (1984), and Piggott and Sweeney (1985).

[2] Examples include Abuaf and Jorion (1990), Edison (1987), Edison and Klovland (1987), Frankel (1986), Frankel (1989), Froot and Rogoff (1994), Kim (1990), and Lothian and Taylor (1993).

[3] This approach has also been pursued independently on smaller panels of post-Bretton Woods OECD data by Lothian (1994), Wei and Parsley (1995), Wu (1994); see also Jorion and Sweeney (1994) and Taylor (1988). Reassuringly, all three studies find strong evidence that PPP tendencies can be found with panel data, a result consistent with our paper. Lothian uses both of our econometric techniques on 22 partners of the United States during the recent dollar float; Wei and Parsley use more dis-aggregated data for 91 country pairs during the post-1973 era, and find that the half-life of PPP deviations is between four and five years; Wu also rejects a unit-root in the real exchange rate on the basis of a panel of recent data. MacDonald (1995), Oh (1995), Papell (1995), and Pedroni (1995) also provide recent papers with similar and complimentary approaches and findings.

Our panel and cross-sectional estimates turn out to be similar to those found in long time-series data. Our favorite point-estimate of the degree of mean-reversion in the annual real exchange rate is 0.85, so that fifteen percent ($1-0.85$, converted to a percentage) of PPP deviations are eroded annually. These estimates are statistically significant, and consistent with the existing time-series literature: raising 0.85 to the fourth power shows that our estimate implies that half of a PPP deviation is closed after four years, the same estimate found with time-series techniques. Thus, our findings should be viewed as complementary to and consistent with those of the existing literature. However, cross-sectional data appear to give more powerful evidence of long-run PPP than do time-series. Observations at a typical point in time across countries appear to be "more independent" and certainly have more variation than do observations for a typical pair of countries over time.

2. Methodology

Our purpose is to compare panel and cross-sectional results with those derived from time-series. To facilitate this comparison, we begin with a standard equation. We estimate:

$$\Delta s_{it} = \alpha + \beta(\Delta p - \Delta p^*)_{it} + \{\Sigma_i \delta_i D_i\} + \{\Sigma_t \phi_t D_t\} + \epsilon_{it} \quad (1)$$

where: Δ denotes the first-difference operator; i denotes country, and t denotes year; s denotes the natural logarithm of the number of units of currency needed to purchase one American dollar; p (p^*) denotes the natural log of the domestic (American) CPI; D_i (D_t) denotes a country-specific (year-specific) "fixed effect" dummy variable intercept; and ϵ denotes a stationary disturbance term. Only the stationarity of the latter is necessary, since it would be unreasonable to think of PPP as holding continuously. Thus, ϵ may be highly autoregressive, representing transitory departures from PPP or some other disequilibrium dynamics. (Throughout, we think of Eq. (1) as being a non-structural linear projection.) We think this strategy is reasonable, since we are only interested in modelling the longer-run tendencies of the data, and in particular the tendencies (or lack thereof) towards PPP, rather than a more complete characterization of the data including the short-run dynamics.

The coefficient of interest to us is β. A finding that β is statistically indistinguishable from unity constitutes confirmation of PPP (technically speaking "relative" PPP, since the equation is estimated in first-differences of logs).[4] On the other hand, it is not clear what alternative interpretation can be given if β is estimated to be different from one.[5]

We follow the literature in estimating Eq. (1) with ordinary least squares. OLS estimates of β are consistent under the hypothesis that PPP deviations are uncorrelated with inflation differences. While it is traditional to make such assumptions, they may be implausible. For this reason, we also estimate (1) with instrumental variables, using a time trend and a single lag of both Δs and $(\Delta p - \Delta p^*)$ as instrumental variables (following the discussion in Froot and Rogoff).

Eq. (1) models the percentage change in the exchange rate as a function of the inflation differential. While informally this direction of causality seems appropriate for countries with floating exchange rates, many countries have fixed their exchange rates for at least part of our sample. In any case, the regression specification is ad hoc, if standard, leaving unresolved potentially important issues of endogeneity, as well as the issue of assuming orthogonality of PPP disturbances to inflation differentials. Thus, we also run the "reverse" regression to Eq. (1), projecting inflation differentials on exchange rate percentage changes. The potential presence of heteroskedasticity leads us to estimate our coefficient covariance matrix with a White/Huber estimator throughout.

We have no strong prior views about the relevance of country- or time-specific fixed effect terms. We check to ensure that our results are insensitive to their inclusion. Indeed, we perform a number of robustness checks on Eq. (1). We estimate it: on only post-1973 data; on only data for industrialized countries; on data averaged over a number of years; and on observations with only small or large values of the inflation differential (so as to keep track of the relative importance of outlier observations).

We will also provide more direct evidence on mean-reversion in the real exchange rate by estimating the following equation:

$$\Delta q_{it} = \alpha + \gamma q_{it-1} + \{\Sigma_i \delta_i D_i\} + \{\Sigma_t \delta_t D_t\} + \epsilon'_{it} \tag{2}$$

[4] As shown by Taylor (1988), the hypothesis of $\beta = 1$ is only truly relevant in the absence of artificial (protectionist) and natural (transportation) barriers to trade, and measurement error in prices.

[5] For the reasons given in Davutyan and Pippenger (1985), Krugman (1978) and Frankel (1986), Frankel (1989), we believe that this equation may not be especially revealing. Essentially, under the null hypothesis—that PPP holds except for random deviations that are small and transitory—it relies on the assumption that PPP deviations are uncorrelated with inflation rates, while it does not make sense at all under the alternative. However, we begin with Eq. (1) to facilitate comparison with the literature.

where $q \equiv s - (p - p^*)$ denotes the natural logarithm of the real exchange rate.

This framework is close in spirit to a traditional time-series Dickey–Fuller test of the proposition that the real exchange rate follows a martingale. Significant negative estimates of γ would indicate substantial mean-reversion in the real exchange rate. The limiting case of $\gamma = -1$ represents complete mean reversion (within the year); $\gamma = 0$ represents no mean reversion, so that the real exchange rate follows a random walk.

However, the panel nature of our set-up means that traditional Dickey–Fuller critical values are inapplicable to test the null hypothesis H_0: $\gamma = 0$. Quah (1994) shows that the relevant critical values for "*t*-like" hypothesis tests concerning γ are quite close to normal for our sample, when all intercepts are suppressed. Levin and Lin (1992) generalize Quah's analysis. They show that the critical values which are appropriate in the presence of a single intercept are nearly normal for our sample. However, they also find that inclusion of a set of country-specific intercepts drives the critical values required to reject the hypothesis H_0: $\gamma = 0$ above 10 in absolute value.

We place greater weight on the estimation results from Eq. (2) than those which stem from Eq. (1). The former provides a superior empirical framework for tests of PPP, since it provides a range of well-specified, economically meaningful hypotheses. One interesting null hypothesis is $\gamma = -1$, i.e., the deviations from PPP are purely transitory. Another is $\gamma = 0$, which implies a complete absence of mean-reversion in the real exchange rate. The intermediate values of γ correspond to different speeds of mean-reversion. These statements are not true of Eq. (1), which does not have an economically interesting alternative hypothesis.[6]

3. The data set

Our data set is annual, and was extracted from the 8/93 cd-rom version of the IMF's *International Financial Statistics*. We use the CPI (IFS line 64) as the measure of prices, and the price of an American dollar (IFS line rf) as the exchange rate. Both of these variables are standard choices for the literature. Series are available for 150 countries, though many countries do not have data which span the full data range, 1948 through 1992 (in which case we use whatever data are available). Throughout, the United States is treated as the base country for both prices and exchange rates.[7] Both

[6] Further, the coefficient of interest is consistently estimable under more plausible circumstances in Eq. (2), as noted earlier.

[7] Our results are not changed substantively if Germany is used as the base country.

exchange rates and CPIs are converted by taking first-differences of natural logarithms.[8]

The raw data set is presented graphically in Fig. 1. Each of the nine "small multiple" graphic images is a scatter-plot of the first-difference in the exchange rate against the inflation differential. Individual observations are marked with dots; the dots are connected with a non-linear non-parametric data smoother. A number of the scatter plots are bordered by pairs of box-and-whiskers graphs, one for each marginal distribution (inflation differential above, percentage change in the exchange rate to the right). These graphical representations of the marginal distributions enable one to determine the location of tight clusters of data.[9]

The nine graphs represent a number of different cuts at the same data set. The entire panel is portrayed at the extreme top left-hand corner of the

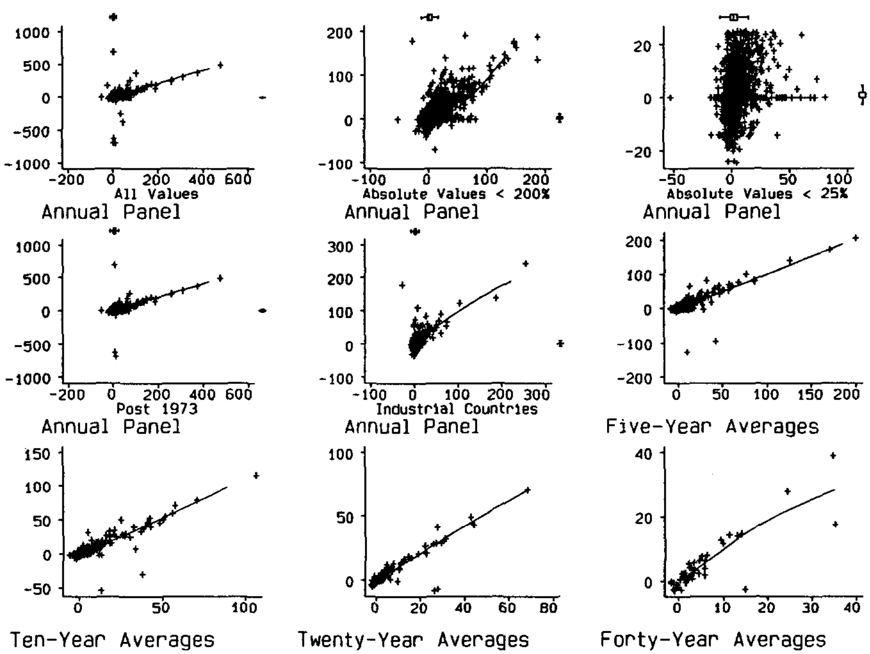

Fig. 1. Raw panel data of exchange rate changes on inflation differentials (1949–1993, all IFS countries, different frequencies, $ rates and CPIs; non-parametric smoothers also shown).

[8] Our STATA data set and programs are available upon receipt of two formatted high-density 3.5 in diskettes and a self-addressed stamped mailer.

[9] The box covers the inter-quartile range with the median marked explicitly inside; the whiskers extend to 150% of the inter-quartile range rolled back to the nearest available observation.

figure. However, most of the observations lie within a small area of this graph, owing to the presence of a few outliers which dominate the plot. To allow one to focus on non-outliers, the other graphs on the top row narrow the range of the data plots by restricting the values of exchange rate percentage change to no more than 200% in absolute value (in the middle) and no more than 25% (on the right). The left and middle graphs in the center row portray post-1973 and industrial country observations only. The last four graphs portray data averaged over five, ten, twenty, and forty years respectively.

Throughout, there is clear evidence of a strong positive correlation between the percentage change in the exchange rate and the inflation differential. This is especially true when high-inflation observations are included.

Exchange rate percentage changes and inflation differentials have similar sample means over the entire sample (6.7% and 5.8% respectively), but very different standard deviations (35.0% and 18.6% respectively). Moreover, this variation differs systematically across the time- and country-dimensions of the panel data set. Table 1 contains some relevant descriptive statistics. It presents one aspect of the relative contributions of time-series variation and cross-sectional variation for the two variables. The top panel of the table presents results computed using only the time-series variation in the data, the average (and standard deviation of this average) across the 150 country-specific time-series, and the average standard deviation for these 150 time-series. The bottom panel is the analog for the 44 year-specific cross-sections. The sample means of exchange rate percentage changes and inflation differentials are quite similar across time and countries.[10] However, the typical standard deviation of the data is much higher (for both the regressor and the regressand of Eq. (1)) across countries than across time. We shall see that the greater variability in the cross-section dimension allows for more powerful tests.

Table 1
Descriptive statistics

	Mean	(Std Dev)	Avg Std Dev
Time-series results for 150 countries			
Exchange rate percentage change	6.8	(11.2)	18.9
Inflation differential	5.7	(9.8)	9.4
Cross-sectional results for 44 years			
Exchange rate percentage change	6.6	(7.2)	27.9
Inflation differential	5.2	(3.6)	14.5

[10] They differ slightly because of the imbalanced nature of the panel; not all countries have observations for all time periods.

4. Results

Table 2 contains estimates of Eq. (1). There are three different panels in the table, respectively referring to: benchmark OLS estimates of Eq. (1) at the top; instrumental variable estimates in the middle; and reverse regressions at the bottom. The different rows correspond to different perturbations of the specification, e.g., including country- or year-specific dummies (i.e., the $\{D_i\}$ and $\{D_t\}$ terms), restricting the sample in various ways, and averaging the data over four different time horizons. The "slope" tabulated is the point estimate of β. The (heteroskedasticity-consistent) standard error is recorded in parentheses. Also tabulated is the sample size N, the R^2 of the regression, and an estimate σ of the root-mean squared error of the residual ϵ.

Table 2
Estimations of Eq. (1), and perturbations thereof

Regressions of percentage change in exchange rate on inflation differential					
	Slope	(se)	N	R^2	σ
Whole panel	0.97	(0.03)	4109	0.29	28.3
Country dummies	0.95	(0.04)	4109	0.30	28.5
Year dummies	0.97	(0.03)	4109	0.31	28.0
Post-Bretton-Woods	0.99	(0.02)	2268	0.40	28.1
Industrial countries only	0.91	(0.06)	1129	0.48	11.9
$\lvert$Values$\rvert$ < 50%	0.77	(0.08)	4016	0.05	27.7
$\lvert$Values$\rvert$ < 20%	0.63	(0.07)	3798	0.02	26.4
$\lvert$Values$\rvert$ < 10%	0.56	(0.10)	3395	0.01	24.6
$\lvert$Values$\rvert$ > 10%	1.01	(0.04)	714	0.44	41.5
Five-year averages	1.01	(0.04)	733	0.73	9.2
Ten-year averages	1.01	(0.06)	330	0.76	6.8
Twenty-year averages	0.96	(0.07)	140	0.82	4.9
Forty-year averages	0.86	(0.17)	48	0.76	3.9
Instrumental variable regressions (instrumental variables in parentheses)					
IV (time)	0.98	(0.15)	4109	0.29	28.3
IV (lag of Δs and $\Delta(p-p^*)$)	0.99	(0.03)	3975	0.29	28.4
Regressions of inflation differential on percentage change in exchange rate					
Whole panel	0.30	(0.08)	4109	0.29	15.6
Country dummies	0.24	(0.07)	4109	0.42	14.3
Year dummies	0.29	(0.08)	4109	0.32	15.4
$\lvert$Values$\rvert$ < 50%	0.07	(0.02)	4016	0.05	8.3
$\lvert$Values$\rvert$ > 10%	0.44	(0.13)	714	0.44	27.2
$\lvert$Values$\rvert$ > 20%	0.53	(0.12)	311	0.52	33.0
$\lvert$Values$\rvert$ > 50%	0.69	(0.12)	93	0.70	38.1
Post-Bretton-Woods	0.41	(0.12)	2268	0.40	17.9
Industrial countries only	0.53	(0.13)	1129	0.48	9.0

OLS results, heteroskedasticity-consistent standard errors (except for IV results). USA is base country; 1948–1992.

The results from the top panel are consistent with (relative) PPP in the sense that β is typically estimated to be close to unity in economic and statistical terms. (In few cases is β significantly different from zero at traditional confidence levels.) For instance, the top row of Table 2 indicates that estimation of the most naive form of Eq. (1) delivers an estimate of $\beta = 0.97$, essentially indistinguishable from the null hypothesis of $\beta = 1$. This result is also relatively insensitive, for example, to inclusion of the different set of fixed-effect intercepts, to restricting the sample to only post-1973 or industrial country data, and to estimation with instrumental variables. These results are quite consistent with those of Lothian (1994) and Wei and Parsley (1995), who both used post-1973 panels of OECD countries. Consistent with Flood and Taylor (1994), averaging the data over time leads to a tight-fitting proportionate relationship between inflation differentials and the change in the exchange rate. β falls if outliers are excluded, which is intuitively predictable.[11]

The bottom panel of Table 2 indicates that the bivariate correlation between inflation differentials and exchange rate percentage changes remains significantly greater than zero when the reverse regression is estimated. However, since inflation differentials are much less volatile than exchange rate percentage changes, the regression coefficients are much smaller in the reverse regression specifications. Succinctly, results from the reverse regressions are much less supportive of PPP than the standard estimates. The ambiguity of the results derived from Eq. (1) is one of the reasons we go on to provide further evidence on mean-reversion in the real exchange rate, using Eq. (2) as our specification.

It is interesting to compare the time-series and cross-sectional estimates that can be derived from our panel, since the innovation in our study is the addition of the cross-sectional variation. Fig. 2 and Fig. 3 provide some relevant evidence, continuing with the reverse regression as the default specification. We estimated Eq. (1) across countries for a given year; Fig. 2 is a histogram of the point estimates of β which are found from the 44 different cross-sections in the data set. Fig. 3 is an analogous histogram of β estimates derived using the 150 country-specific time-series in the data. The central tendency in β derived from the cross-sectional data is both higher and more precise than the central tendency derived only from time-series variation.

Table 3 contains estimation results for Eq. (2), the equation which estimates the mean-reversion in the real exchange rate. The three different panels in the table correspond to three different assumptions about the

[11] The top right-hand graph in Fig. 1 shows clearly that while the volatility of the regressand is restricted, there is no comparable restriction on the range of the regressor, leading to a lower estimate of β.

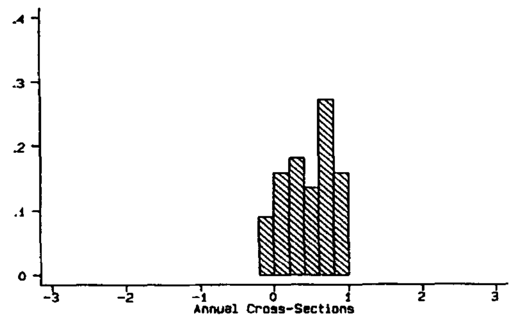

Fig. 2. Cross-section estimates of β (PPP slope regressions; mean = 0.45; standard deviation = 0.33).

intercepts in Eq. (2). The top panel suppresses both country- and year-specific intercepts; the middle panel adds country-specific intercepts; the bottom panel, year-specific intercepts. Each specification is estimated with a variety of restrictions on the data set.[12,13]

In all cases, the point estimate of γ is negative. The central tendency is around -0.15, which implies a half-life of around four years, consistent with the existing time-series literature. It is somewhat more difficult to establish the statistical significance of these estimates, because of the

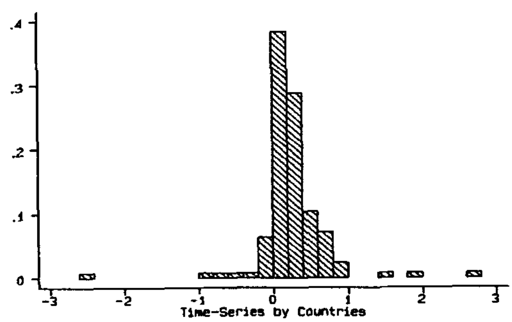

Fig. 3. Time-series estimates of β (PPP slope regressions; mean = 0.25; satndard deviation = 0.48).

[12] We normalize the real exchange rate by expressing it as the deviation from the country-specific mean.

[13] Table 3 reports heteroskedasticity-consistent standard errors, so as to provide precise confidence for our point-estimates of γ. Traditional OLS standard errors, which would be appropriate for Dickey–Fuller tests of unit-root non-stationarity, are even smaller, further strengthening our results. Also, our results are insensitive to inclusion of either one or two lags of the regressand, as would be appropriate in an augmented Dickey–Fuller set-up; the results for the case of a single augmenting lag are shown in Table 3.

Table 3
Estimates of Eq. (2), and perturbations thereof

First-difference of log real exchange rate on lag log real exchange rate					
No fixed effects	Slope	(se)	N	R^2	σ
Whole panel	−0.12	(0.04)	4109	0.06	0.274
|Values| < 50%	−0.01	(0.004)	4060	0.00	0.098
|Values| > 10%	−0.28	(0.09)	923	0.14	0.549
With lagged difference	−0.10	(0.04)	3975	0.09	0.271
Post-Bretton-Woods	−0.12	(0.06)	2268	0.05	0.274
Industrial countries	−0.15	(0.06)	1129	0.09	0.114
Country-specific intercepts					
Whole panel	−0.12	(0.04)	4109	0.07	0.276
|Values| < 50%	−0.01	(0.004)	4060	0.04	0.098
|Values| > 10%	−0.32	(0.10)	923	0.20	0.569
With lagged difference	−0.10	(0.04)	3975	0.11	0.273
Post-Bretton-Woods	−0.40	(0.14)	2268	0.25	0.250
Industrial countries	−0.15	(0.06)	1129	0.11	0.114
Year-specific intercepts					
Whole panel	−0.12	(0.04)	4109	0.09	0.271
|Values| < 50%	−0.01	(0.005)	4060	0.18	0.090
|Values| > 10%	−0.28	(0.08)	923	0.24	0.530
With lagged difference	−0.10	(0.04)	3975	0.12	0.268
Post-Bretton-Woods	−0.12	(0.06)	2268	0.08	0.270
Industrial countries	−0.18	(0.08)	1129	0.45	0.091

OLS results, Huber/White standard errors. USA is base country; 1948–1992.

complications associated with unit-root tests and panel data sets. Levin and Lin (1992) show that the 5% critical value for "t-like" tests of H_0: $\gamma = 0$ is around -1.8 when only a single intercept is included; the 1% value is around -2.4. Most of our t-tests in the top panel exceed these values comfortably. In other words, they reject the hypothesis that the real exchange rate follows a random walk. But while our t-ratios do not change much with the inclusion of country-specific fixed effects, the relevant critical values jump enormously, as shown by Levin and Lin. Thus, we are not able to reject the null hypothesis of H_0: $\gamma = 0$ when country-specific intercepts are included. On the other hand, neither the fit nor the slope estimates of these regressions seems much affected by the inclusion of country intercepts.[14] It is also interesting to note that point-estimates from data which use only floating-rate observations are similar to those also use data from regimes of fixed exchange rates. While mixing data from different exchange rate regimes is potentially inappropriate, it turns out not to be an issue in this sample.

Mean reversion in the real exchange rate can be seen in Fig. 4, a graphical

[14] Levin and Lin do not address the case of time-specific intercepts.

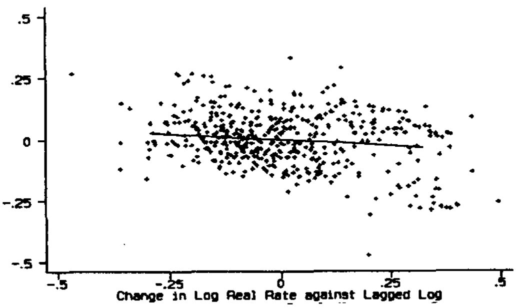

Fig. 4. Mean-reversion in the real exchange rate (industrial countries, 1974–1992; slope estimate = −0.18 (standard error = 0.03)).

version of (one estimate of) Eq. (2). Fig. 4 is a scatter plot of the change in the log of the real exchange rate against its lagged level (after allowance for country-specific intercepts); a non-parametric data smoother is also included to help "connect the dots". Evidence of mean-reversion is apparent. For the sake of variety, only post-1973 industrial country data are plotted in Fig. 4. The slope estimate of γ is −0.18 with a standard error estimated to be 0.03 (when country-specific intercepts are included).[15]

Next we test pure cross-section and time-series versions of this equation. We perform our analysis for all 19 post-1973 year-specific cross-sections, and for all 131 post-1973 country-specific time-series.

For seventeen of the nineteen years, γ is estimated to be negative. It is significantly different from zero at the 1% level in five of these regressions. By way of contrast, only eight of the γ estimates are significantly negative (at the 1% level) for the 131 time-series. Histograms of the t-statistics are displayed in Fig. 5 and Fig. 6 for the 19 cross-sections and 131 time-series respectively.

We have gone to lengths to show that our results in support of PPP and reasonable mean-reversion of the real exchange rate are robust to a variety of modifications of our basic empirical methodology. We have also performed, but not reported, a number of additional checks, mostly combinations of our various restrictions. Our results appear to be quite insensitive.[16]

[15] When Eq. (2) is estimated on this post-1973 industrialized country panel (with country-specific intercepts), the t-test for H_o: $\gamma = 0$ exceeds 5. However, when Eq. (2) is estimated on a country-by-country basis using only the time-series variation in the data, none of the t-tests for $\gamma = 0$ is significantly different from zero at the 5% significance level.

[16] We have also found, in tests not reported, that the real exchange rate tends to revert towards *absolute* PPP, using the Penn World Table measure of absolute PPP deviations.

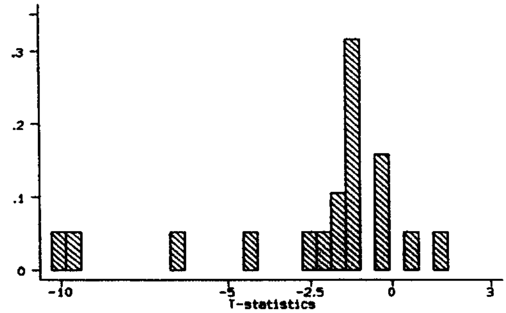

Fig. 5. Cross-sectional t-statistics (mean = 2.4; standard deviation = 3.1).

5. Why is variation in the data so important?

The analysis we have presented shows that it is relatively simple to find support for PPP using panel data. This conclusion is at odds with the conventional view that it is not easy to reject the hypothesis that the real exchange rate follows a random walk. In this section, we show how it is possible to reconcile our results with those from the time-series literature, once the variation of the data is examined in detail.[17]

Support for PPP derived from Eq. (1) consists of a *failure to reject* the null hypothesis H_0: $\beta = 1$. As stated above, it is hard to believe that PPP deviations (the ϵ terms in Eq. (1)) are completely uncorrelated with inflation differentials. If PPP deviations are correlated with inflation differentials, then OLS estimates of β are biased and inconsistent because of the errors-in-variables problem. However, the null hypothesis of a unit coefficient will emerge, as the size of the PPP deviations becomes suffi-

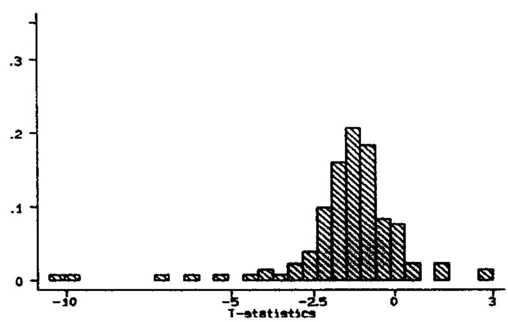

Fig. 6. Time-series t-statistics (mean = 1.4; standard deviation = 1.7).

[17] For a complementary approach, see Edison et al. (1994).

ciently small relative to the total variation in the data. The reason is that the bias vanishes as data variation rises; see Davutyan and Pippenger (1985). As the descriptive statistics in Table 1 show, cross-sectional variation is higher than time-series variation. Panel OLS estimation of β thus has two advantages over pure time-series regressions: more volatility, and more data.

Consider next tests of Eq. (2). In this case, support for PPP consists of *rejecting* the null hypothesis of no mean-reversion in the real exchange rate (i.e., rejecting the random walk hypothesis). It is easy to show that a data set with insufficient total variation may fail to reject the null hypothesis because of inadequate power. A pure time-series variant of Eq. (2) can be re-written as a simple autoregression, $q_t = \phi q_{t-1} + \epsilon_t$ (ignoring any intercept). The asymptotic standard error of an estimate of ϕ is approximately the square root of $(1-\phi^2)/N$. If the true speed of adjustment is 15 per cent a year ($\phi = 0.85$), a simple calculation suggests that we might require more than a century of data to be able to reject the null hypothesis H_o: $\phi = 1$ using a time-series approach.[18] It is not very surprising that 45 years of data is not enough, much less the 20 years of data available since 1973.[19] Again, the message is that the volatile large panel allows one to estimate mean-reversion tendencies with greater precision than short time-series regressions.

6. Conclusions

This paper examines purchasing power parity using a panel data set of 150 countries and 45 annual observations. Our results are consistent with the emerging consensus view that deviations from PPP have a half life of approximately four years. It is difficult to find such results with a pure time-series approach; one is forced to rely on a century of data which is frequently unavailable. Long samples also necessitate pooling data across exchange rate regimes which is potentially troubling (although, as it turns out, this makes no difference in our sample). Still, it is much easier to find the requisite variation in the data by exploiting cross-sectional variation.

[18] The reason is that $2.93^2(1-.85^2)/(1-.85)^2 = 106$. Further, the gain in power from using a higher frequency data set is small, as shown in Frankel (1986), Frankel (1989).

[19] Econometricians consider the asymptotic standard error on which this calculation is based to be a bad approximation in small samples. But the correct power calculation suggests that, if anything, the sample required to reject a random walk would be even larger than 106. DeJong et al., (1988, Table 2) offer power tables for the Dickey–Fuller test which show that when the true ϕ parameter is 0.8, even a sample size of 100 is sufficient to reject a random walk only about 65 per cent of the time.

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