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Multi-country evidence on the behavior of purchasing power parity under the current float

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Using panel data for the United States and 22 other OECD countries for the current float, this paper presents evidence that despite substantial short-term perturbations, purchasing power parity actually performed much better than commonly believed. Average rates of growth of real exchange rates over long horizons bear little relationship either to average rates of growth of nominal exchange rates or to average inflation differentials, thus implying a close to one-to-one relation between average rates of growth of nominal exchange rates and average inflation differentials. Panel-data variants of standard unit-root tests suggest that the real exchange rates of these countries can be characterized as mean-reverting. (JEL F31). © 1997 Elsevier Science Ltd

In the two decades following the introduction of floating exchange rates, the concept of purchasing power parity fell into increasing disfavor. Nominal exchange rates had tended to move roughly in line with real exchange rates in conspicuous violation of short-term purchasing power parity. Real exchange rates, in the view of many researchers, were characterized by sizable, seemingly permanent shifts, which, if actually the case, would imply long-term violation too. The result until quite recently was near total skepticism about the merits of purchasing power parity both as a theoretical building block and as an empirical rule of thumb (see Dornbusch, 1987; Frankel and Meese, 1987; Meese, 1990).

This paper reexamines that experience. Using data for the United States and 22 other OECD countries during the first two decades of the float, I present

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evidence that despite substantial short-term perturbations PPP actually performed much better than commonly believed. Average rates of growth of real exchange rates over long horizons bore little relationship either to average rates of growth of nominal exchange rates or to average inflation differentials. Translated in terms of purchasing power parity, there was therefore a close to one-to-one relation between average rates of growth of nominal exchange rates and average inflation differentials. Further support for PPP in these data comes from panel-data variants of standard unit-root tests, which suggest that the real exchange rates of these countries can be characterized as mean-reverting.

These findings, therefore, go at least part of the way toward solving an important empirical puzzle—how to reconcile exchange-rate behavior under the float with what has been learned from recent studies of longer term historical data. For the float, researchers typically have concluded that real exchange rates were well approximated as random walks (e.g. Roll, 1979; Darby, 1983). Analyses of long-term historical data, in contrast, have pointed to mean-reverting behavior of one sort or another (e.g. Lothian, 1990; Diebold *et al.*, 1991; Johnson, 1993; and Lothian and Taylor, 1996).

On one view, this difference in results across the two bodies of data is an econometric problem, resulting from the low power of conventional tests in distinguishing between unit-root and near-unit-root behavior in samples that, like the float, span a relatively short number of years (e.g. Frankel, 1986; Lothian, 1990; Edison *et al.*, 1996; Lothian and Taylor, 1997). On an alternative view, the difference is 'real', in both the popular and the economic senses of the word. According to this line of reasoning, the difference in the performance of purchasing power parity under the float and under earlier regimes is not a statistical artifact, but a behavioral difference, a reflection of the greater incidence and severity of real shocks over the past two decades than previously (Stockman, 1990; Grilli and Kaminsky, 1991).

The results presented in this paper lend little support to this latter explanation. Behavior under the float, in fact, appears very similar to that observed in the earlier historical data, suggesting that as a long-run proposition purchasing power parity remained a quite useful first approximation.

I. Theory and recent evidence

According to the purchasing power parity theorem the logarithm of the nominal exchange rate, the foreign-currency price of a unit of domestic currency, e, will equal the difference in the logarithms of the foreign and domestic price levels, $p - p^*$:

$$e_t = p_t - p. \qquad \langle 1 \rangle$$

In growth-rate form, therefore,

$$\hat{e}_t = \hat{p}_t - \hat{p}, \qquad \langle 2 \rangle$$

where a carat over a variable represents its time derivative.

A variety of theoretical models, ranging from simple open-economy versions of the quantity theory of money to the two-country, cash-in-advance model of

Lucas (1982), gives rise to the purchasing-power-parity theorem. All see PPP as an equilibrium position that follows from the homogeneity postulate. In these basic models, PPP holds continuously. In other, more complex models, additional factors are admitted to allow for both short-run and long-run departures. In one such class of models, monetary shocks are posited to be the major source of disturbances (Dornbusch, 1976; Frenkel, 1981). Given price stickiness, such shocks result in transient divergences of nominal exchange rates from PPP, and, as a result, of actual real exchange rates from their equilibrium values. Another major class of models regards disturbances as predominantly real in origin and permanent in their impact (e.g. Stockman, 1980). The terms of trade, productivity shocks and differential rates of productivity growth in tradable-goods sectors are factors that have been cited as sources of permanent divergences between nominal exchange rates and PPP and, hence, of permanent shifts in the levels of real exchange rates. Viewed in terms of these more complex models, purchasing power parity remains a useful theoretical building block, but its empirical importance is an open question.

I.A. Previous evidence from the float

In the early 1980s, monetary and portfolio balance models of exchange rates appeared to break down, producing parameter estimates that were inconsistent with theory and explained little of the variation in nominal exchange rates (Frankel, 1984). As forecasting tools, they proved inferior to simple random-walk models in *ex post* dynamic simulations (Meese and Rogoff, 1983). Time-series models, moreover, indicated that real exchange rates were approximately random walks, and subsequent tests showed that it was impossible to reject either the hypothesis of a unit root in real exchange rates or the less restrictive hypothesis of non-cointegration of their nominal-exchange-rate and relative-price-level components (e.g. Enders, 1988; Taylor, 1988).

Perhaps the most important source of dissatisfaction with PPP under the float, however, was the evidence provided by actual day-to-day observation. Nominal exchange rates—particularly US dollar rates during the early and mid-1980s—showed considerable variation. Much of this variation took the form of protracted swings, swings that in turn appeared to bear little or no, and at times even a perverse, relationship to movements in fundamental economic variables. Indeed, one of the stylized facts of exchange-rate behavior during these years was the substantial correlation between quarter-to-quarter and year-to-year changes in nominal and real exchange rates alluded to above.

As researchers have pointed out, however, the econometric tests, such as unit root tests, can be misleading in small samples.[1] Given the relatively short span of the data for the float, these tests are unlikely to be powerful enough to distinguish between unit-root and near-unit-root behavior. One obvious way to circumvent such data limitations is to use a long historical data set. And in the past several years a considerable number of studies have done exactly that. The evidence emerging from these studies has been much more favorable to purchasing power parity. The common conclusions of this research are that real exchange rates contain sizable mean-reverting components, but that this process of mean reversion is quite slow.[2] Deviations from PPP are therefore persistent, but in the end largely disappear. Such findings, however, have not gained uncritical acceptance.[3] In any event, they stem mainly from investigations of very long time series, in most instances series that span a century or more. Hence, it is quite possible that a structural change did occur following the breakdown of the Bretton–Woods system but that the aggregation of data for the float with the earlier, much longer body of historical data is masking that change.[4]

II. Data and empirical results

The key to being able to distinguish between these two competing explanations for why results differ so markedly across data sets is experimental design. One possible approach would be to use long-term historical data, and to focus on the stability of real exchange rates across regimes, rather than on their time-series behavior *per se*.[5] Another would be to use data for the float alone and to estimate directly the permanent and transitory components of real exchange rates (e.g. Evans and Lothian, 1993). A third would be to pool the floating-rate data for a number of countries, and to conduct tests using this pooled sample.[6]

In this paper I adopt a variant of this last approach. I use multi-country time series data, but as in Lucas (1980), Duck (1993) and Lothian (1985), I concentrate on long-run behavior within these countries, and exploit the cross-sectional aspects of the data to test and otherwise evaluate the relationships of interest. One advantage of this approach is its simplicity. The graphical evidence presented below is both transparent and easily replicatable. A second advantage is that in analysing the differences in behavior among countries it utilizes a potentially quite rich body of information that typically is ignored.[7] A disadvantage is that in averaging and taking rates of growth, some of the information in the time series is lost. For that reason I also conduct a set of unit root tests of the (log) levels of the real exchange rates using the full panel data set.

II.A. An overview of the data

The sample used in this investigation encompasses the United States and 22 other OECD countries over the period 1974 through 1990. Exchange rates are denominated in US dollars; the price-level measure is the consumer price index, or cost-of-living index, depending upon the country.[8]

The problems with purchasing power parity that I have described for the major currencies are also characteristic of the data for this broader sample of countries. We can see this quite clearly in Figure 1 and in the results of the regressions and associated unit-root tests reported in Table 1. Plotted in Figure 1 is the 22-country average normalized log-real dollar exchange rate for the years 1974–1990. Of particular interest, in this graph, are the volatility of this index and the similarity of its path to those followed by the more familiar (weighted) real-exchange-rate indexes of the International Monetary Fund and

TABLE 1. DF tests for average and individual real dollar exchange rates of 22 OECD countries, 1974–1990, $\Delta q_t = \mu + \lambda q_{t-1} + \epsilon_t$

	λ	DF		λ	DF
Average	– 0.19	–1.91	Italy	–0.27	–2.02
Australia	–0.22	–2.03	Japan	–0.08	–1.08
Austria	–0.18	–1.81	The Netherlands	–0.20	–2.04
Belgium	–0.24	–2.32	New Zealand	–0.27	–2.20
Canada	–0.19	–2.03	Norway	–0.14	–1.56
Denmark	–0.18	–1.81	Portugal	–0.15	–1.48
Finland	–0.13	–1.26	Spain	–0.18	–1.76
France	–0.28	–2.55	Sweden	–0.24	–2.23
Germany	–0.22	–2.13	Switzerland	–0.13	–1.49
Greece	–0.25	–2.29	Turkey	–0.12	–1.47
Iceland	–0.27	–2.33	United Kingdom	–0.38	–3.05
Ireland	–0.27	–2.16			

Note: DF is the Dickey–Fuller test statistic.

the Federal Reserve Board. Since these latter series, in turn, have behaved very similarly to the real dollar exchange rates of the major currencies, the currencies that have been the subject of most empirical investigations, the sample used here appears to be entirely representative of what has come to be regarded as typical of experience under floating rates.

The regression results reported in Table 1 confirm this impression. For both the index itself and for the 22 individual countries' real exchange rates, I

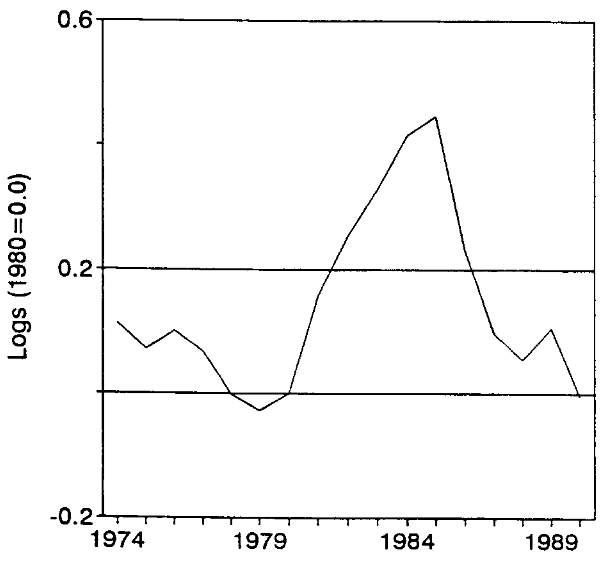

FIGURE 1. Index of real dollar exchange rates of 22 OECD countries.

estimated autoregressions of the form

$$\langle 3\rangle \qquad \Delta q_t = \mu + \lambda q_{t-1} + \epsilon_t,$$

and conducted Dickey–Fuller unit-root tests of the hypothesis $\lambda = 0$. These tests produced results very similar to those obtained in other studies of this period: in each instance, λ was less than zero, but in every case, except that of the United Kingdom, the difference was not statistically significant.

The other important item to note in Figure 1 is the pattern of fluctuations in the index, the dominant roles of a few substantial and protracted movements in the series in accounting for its overall variance, and the tendency for such movements to offset one another through time. We see a large increase and then decrease in the index in the early and mid-1980s, respectively; a diminution in the amplitude of fluctuations thereafter; and, on net, an approximate canceling out over the period as a whole. Visually, there is the suggestion of mean reversion, albeit rather slow mean reversion.

Again, these are features shared by both the published indexes and the individual real dollar exchange rates of the major currencies. But because the two movements that I have pointed to have dominated all of the series, and because the sample period is relatively short to begin with, it is not at all surprising that most researchers have been unable to reject the hypothesis of a unit root in the real exchange rate for the float.

II.B. Cross-country results

The question of real-exchange-rate behavior under the float therefore depends importantly on the nature of the two major movements in real exchange rates that we see in Figure 1, whether the approximate offset that we see there is a behavioral phenomenon, as PPP would suggest, or simply the result of happenstance.[9] The added degrees of freedom in cross-country data allow us to speak directly to this issue. Figures 2–4 and the related regressions reported in Table 2 summarize the principal results of this analysis. In all instances, the data are in the form of period averages for the various countries over the years 1974–1990.

Figures 2 and 3 plot the average rates of growth of the 22 countries' real dollar exchange rates relative to their two additive components, the inflation differential in the case of Figure 2 and the average rate of growth of the nominal exchange rate in the case of Figure 3. Figure 4 plots the one component relative to the other and hence is a linear transformation of the relation plotted in Figure 2. I include it to provide a more direct view of PPP performance.

In the first two charts, the picture is virtually the same. In both instances, the individual points appear to be scattered fairly closely about a horizontal line, which, if actually plotted, would be not far removed from the horizontal axis. There is, moreover, no apparent difference in either of the scatters at high versus low values of $\hat{e}$ or $\hat{p} - \hat{p}^*$. What is true for the average real exchange rate plotted in Figure 1 is therefore also true for the individual countries' real exchange rates: over the long run, we see very little net change.

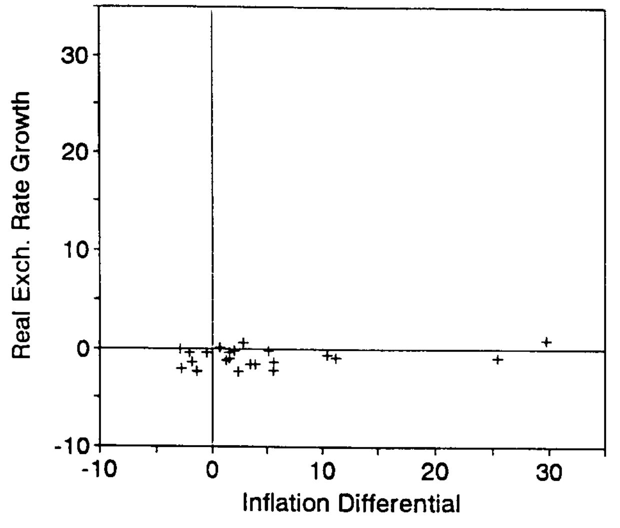

FIGURE 2. Real exchange rate growth and inflation differentials.

Figure 4, which plots $\hat{e}$ relative to $\hat{p} - \hat{p}^*$, is particularly striking. All of the points in the chart are quite tightly clustered about a 45-degree line drawn through the means. Hence, as suggested by the two previous charts, there is virtually a one-to-one relationship between the movements in the two variables.

The regressions reported in Table 2 add numerical precision to these visual impressions. In the first two regressions, $\hat{q}$ is the dependent variable. In these regressions, the slope coefficients are close to zero, and in both instances less

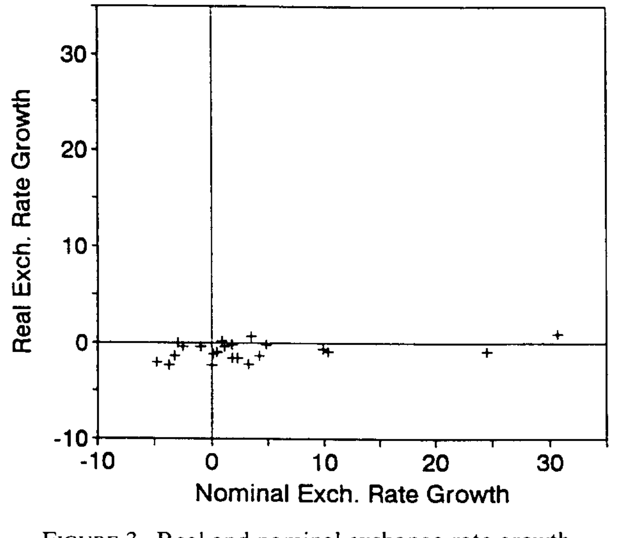

FIGURE 3. Real and nominal exchange rate growth.

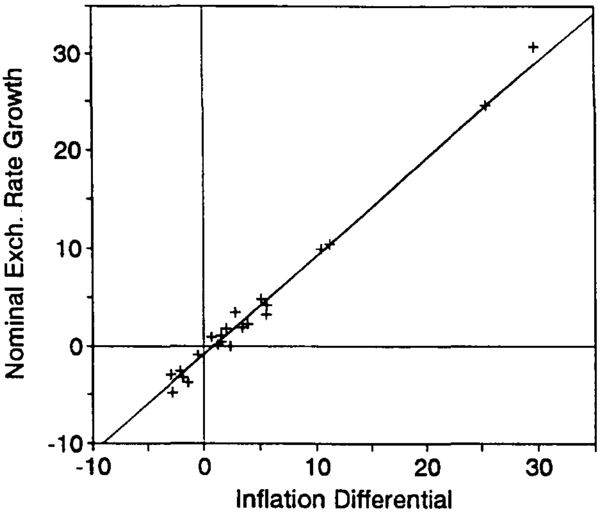

FIGURE 4. Nominal exchange rate growth and inflation differentials.

than twice their estimated standard errors. In the latter two regressions, $\hat{e}$ and $\hat{p}-\hat{p}^*$ alternate as dependent and independent variables. These are the purchasing-power-parity analogues of the real-rate regressions just reviewed. The slope coefficients in these regressions, as one would expect given that they are simply linear transformations of the previous two, are almost identically unity.[10] They are, moreover, over forty times their respective standard errors.

The intercepts of the regressions, however, are non-zero and statistically significant. Contrary to the implications of equation ⟨2⟩, the rate of growth of the nominal exchange rate on average differed from the inflation differential.

TABLE 2. Cross-country regressions: annual averages for 22 OECD countries for the years 1974–1990

Dependent variable	Coefficients and summary statistics				
	Intercept	$\hat{e}$	$\hat{p}-\hat{p}^*$	R^2	SEE
$\hat{q}$	−0.011	0.043		0.160	0.009
	−5.16	1.95			
$\hat{q}$	−0.011		0.034	0.092	0.009
	−4.70		1.43		
$\hat{e}$	−0.011		1.034	0.989	0.009
	−4.70		43.40		
$\hat{p}-\hat{p}^*$	0.011	0.957		0.989	0.009
	5.16	43.40			

Note: Figures beneath the coefficients are t statistics.

Nevertheless, the disparity between the two is small, both in absolute value and in comparison to the observed movements in real exchange rates over the period. As an example, consider the index of real dollar exchange rates plotted in Figure 1. The cumulative increase in the index from 1980 to 1985 was 45% at a continuously compounded rate, an average pace of 9% per annum over these 5 years; the cumulative decrease from 1985 to 1988 was 39%, an average pace of 13% per annum. The estimated pairs of intercepts converted from decimal to percent terms, in contrast, are both 1.1% per annum. The standard errors of estimate of the respective regressions, when compared with the standard deviations of the variables themselves, tell a similar tale. Both are less than 1% per annum (0.88 and 0.91) as opposed to a standard deviation of the average rate of change of the nominal exchange rate of 8.8% per annum and a standard deviation of the average inflation differential of 8.4% per annum. The R^2 in the two PPP variants is therefore extremely high—0.989. More importantly, it remains high (0.977) if we impose the theoretical constraint of a zero intercept and a unit slope coefficient.

So, while PPP may not have held rigidly during this period, even over the long run, the results that we have just reviewed suggest that as a first approximation it nevertheless performed quite well. The additional evidence presented immediately below supports this conclusion. It speaks to three related questions: the time horizon over which approximate convergence between nominal exchange rate changes and inflation differentials occurs; the robustness of the basic results with regard to variations in the sample; and whether convergence between the *levels* of nominal exchange rates and relative price levels can be detected in these data.

II.C. Long-run vs short-run behavior

Tables 3 and 4 present the evidence on long-run vs short-run behavior, and relatedly on the length of time it takes for long-run behavior to emerge in the data. Table 3 contains the results of cross-country regressions of $\hat{q}$ on $\hat{e}$ for various levels of aggregation of the data, ranging from none at all—regressions for the individual years—to averages taken over 3-year, 6-year and 9-year subperiods, as well as those for the full period that were presented above. These figures show a more or less gradual progression from the often strong relationships between the two variables in the yearly regressions to weak and insignificant relationships in the regressions with the 6-year averages, to even weaker relationships in the regressions with the 9-year and the full-period averages. Viewed on the basis of these results, the long run appears to be somewhere between 3 and 6 years, an estimate that is consistent with the estimated half lives of deviations from long-term equilibrium of 2–5 years found in various time-series analyses of real exchange rates (e.g Abuaf and Jorion, 1990; Lothian and Taylor, 1996).

Shown in Table 4 are the results of three regressions run on the pooled annual time series for the 22 countries combined. These highlight further the difference between the results for the full-period-average and the year-to-year relationships between real and nominal exchange rates. The first regression,

TABLE 3. Cross-country regressions of percentage change in real exchange rate on percentage change in nominal exchange rate: yearly and temporally averaged data, $\hat{q} = a + b\,\hat{e} + \epsilon$

	a/Sa	b/Sb	R^2/SEE		a/Sa	b/Sb	R^2/SEE
1974	−0.034	0.060	0.004	1989	0.053	0.004	0.000
	0.011	0.221	0.049		0.017	0.136	0.058
1975	−0.047	0.476	0.689	1990	−0.095	0.209	0.082
	0.008	0.071	0.038		0.019	0.156	0.068
1976	−0.013	0.477	0.548				
	0.012	0.097	0.036	1974–1976	−0.022	0.182	0.192
					0.014	0.084	0.026
1977	−0.043	0.328	0.311				
	0.009	0.109	0.041	1977–1979	−0.042	0.173	0.195
					0.013	0.079	0.035
1978	−0.056	0.401	0.681				
	0.009	0.062	0.041	1980–1982	0.063	0.216	0.311
					0.011	0.072	0.040
1979	−0.024	0.112	0.035				
	0.013	0.131	0.061	1983–1985	0.055	0.073	0.113
					0.009	0.045	0.022
1980	0.011	0.272	0.515				
	0.012	0.059	0.055	1986–1988	−0.109	0.245	0.559
					0.007	0.049	0.025
1981	0.066	0.458	0.374				
	0.029	0.133	0.065	1989–1990	−0.029	−0.110	0.038
					0.010	0.123	0.047
1982	0.041	0.349	0.506				
	0.015	0.077	0.040				
				1974–1979	−0.031	0.090	0.101
1983	0.042	0.203	0.364		0.007	0.060	0.020
	0.013	0.060	0.043				
				1980–1985	0.064	0.116	0.217
1984	0.066	0.148	0.142		0.007	0.049	0.025
	0.014	0.082	0.038				
				1986–1990	−0.087	0.059	0.097
1985	0.019	0.150	0.122		0.004	0.040	0.018
	0.012	0.090	0.039				
1986	−0.122	0.565	0.827	1974–1982	0.009	0.049	0.102
	0.013	0.058	0.040		0.003	0.032	0.012
1987	−0.108	0.263	0.347	1983–1990	−0.033	0.045	0.117
	0.011	0.080	0.034		0.005	0.027	0.012
1988	−0.043	0.116	0.102				
	0.009	0.077	0.044	1974–1990	−0.011	0.043	0.159
					0.002	0.000	0.009

— *continued*

TABLE 3 (*Continued*)

	Summary of regression results							
	a	b	R^2	SEE	a	b	R^2	SEE
Annual					6-Year			
Mean	−0.017	0.270	0.332	0.047	−0.018	0.088	0.138	0.021
Range	0.188	0.562	0.827	0.034	0.151	0.056	0.120	0.007
3-Year					9-Year			
Mean	−0.014	0.130	0.235	0.033	−0.012	0.047	0.109	0.010
Range	0.172	0.355	0.520	0.026	0.042	0.004	0.015	0.003

Note: Sa and Sb denote standard errors of the intercept and slope coefficients, respectively.

which is reported for the sake of completeness, is the simple bivariate pooled regression:

$$\langle 4\rangle \qquad \hat{q}_{it} = \alpha + \beta\hat{e}_{it} + u_{1it},$$

where i is an index of countries.

The second regression is based on the fixed effects model:

$$\langle 5\rangle \qquad \left(\hat{q}_{it} - \bar{\hat{q}}_{i.}\right) = \gamma\left(\hat{e}_{it} - \bar{\hat{e}}_{i.}\right) + u_{2it},$$

where the bar and dot indicate a mean taken over the omitted (time) subscript. This regression includes individual intercepts for countries and thus depicts the (average) time-series relationship between real and nominal exchange rates in the 22 countries.

The third regression has the form:

$$\langle 6\rangle \qquad \hat{q}_{it} = \alpha + \beta_1\hat{e}_{it} + \beta_2\bar{\hat{e}}_{.t} + u_{3it},$$

where $\bar{\hat{e}}_{.t}$ is a vector of the country means. This last equation is a linear

TABLE 4. Pooled regressions: annual panel data for 22 OECD countries, 1974–1990

Dependent variable	Coefficients and summary statistics					
	Intercept	$\hat{e}_{it}$	$(\hat{e}_{it} - \bar{\hat{e}}_{.t})$	$\bar{\hat{e}}_{.t}$	R^2	SEE
$\hat{q}_{it}$	−0.030	0.556			0.599	0.069
	−8.090	23.570				
$(\hat{q}_{it} - \bar{\hat{q}}_{.t})$			0.795		0.841	0.043
			43.100			
$\hat{q}_{it}$	−0.011	0.795		−0.752	0.836	0.044
	−4.240	43.490		−23.180		

Note: See the text for a description of the variables. Figures beneath the coefficients are t statistics.

combination of equation ⟨5⟩ and the (cross-country) regression shown in Table 2, which in the current notation, can be written as:

$$\langle 7\rangle \qquad \bar{\hat{q}}_{.t} = \tilde{\alpha} + \tilde{\gamma}\bar{\hat{e}}_{.t} + \tilde{\epsilon}_t .$$

The coefficient β_2 in ⟨6⟩, which is of principle interest, equals the difference between the slope coefficients in ⟨7⟩ and ⟨5⟩, $\bar{\gamma} - \gamma$; the intercept in ⟨6⟩, α, equals the intercept in ⟨7⟩, $\tilde{\alpha}$. A value of β_2 of zero would mean that the time-series and cross-section coefficients are identical. A test of the hypothesis $\beta_2 = 0$ is therefore a test of the hypothesis that $\gamma = \tilde{\gamma}$. This is the Zellner test for aggregation bias, which is used here to evaluate the homogeneity of the year-to-year and the cross-country relations between $\hat{q}$ and $\hat{e}$. As we can see in the table, changes in real and nominal exchange rates were quite closely related when viewed on a year-to-year basis—an estimated β_1 of 0.795. This stands in rather sharp contrast to the almost complete lack of relationship in the period-average regressions described earlier. The difference between the two is reflected in an estimated β_2 of -0.752. The t-value associated with this estimate is 23.18, leading to an overwhelming rejection of the hypothesis that the two relationships are the same.

II.D. Robustness to variations in the sample

As a check on the robustness of the basic results with regard to variations in the sample, I conducted two additional types of analysis. In the first, I altered the countries included in the full-period-average regressions to see whether the inclusion of countries with very high or very low inflation had unduly influenced the results. In the second, I experimented with data averaged for partially overlapping 6-year and 9-year subperiods to see whether the results were peculiar to the sample period.[11] In both instances, the regressions were of the rate of change of real exchange rates on the rate of change of nominal. In neither case were the estimates substantially different from those reported in Table 2.

Omitting the five lowest and then the four highest inflation countries from the sample resulted in estimated slope coefficients very close to the 0.034 estimate reported in Table 2 for all 22 countries—a value of 0.030 when the lowest inflation countries were excluded, and a value of -0.030 when the highest were excluded. Using the overlapping samples of 6-year-averaged and 9-year-averaged data, I obtained estimates of slope coefficients that again were all quite close to the estimate for the full period. In the case of the 6-year averages, the median of the 12 estimated slope coefficients was 0.063 and the standard deviation of these coefficients was 0.045; in the case of the 9-year averages, the median of the nine estimated coefficients was 0.062, and the standard deviation was 0.029. The similarity between these median estimates and the estimate for the full-period-averaged data reported in Table 2, and the relative homogeneity of the regressions underlying the former suggest that the full-period results were not due to a fortuitous choice of beginning and ending years.

TABLE 5. Results of ADF tests

No. of lags	λ	t-Statistic	Critical values	
			1%	5%
1	−0.29	−9.66	−8.32	−7.69
2	−0.32	−8.98	−8.07	−7.47
3	−0.37	−9.38	−8.02	−7.36
4	−0.46	−11.15	−7.83	−7.28

Note: Critical values are from Oh (1996), Table 2.

II.E. Mean reversion

One way to interpret the results that I have just described is in terms of cointegration. The fact that the fluctuations in real exchange rates in all of these countries dampen appreciably over the full period suggests that the log nominal exchange rates and the log relative price levels of these countries share identical stochastic trends and therefore are cointegrated.

To test this hypothesis, I use the PPP theorem to constrain the cointegrating vector. Specifically, since the log real exchange rate q is defined as

$$\langle 8 \rangle \qquad q \equiv e - p + p^*,$$

I impose the constraint that the cointegrating vector is $(1, -1, 1)$. Given this constraint, the test for cointegration reduces to a test for a unit root in the real exchange rate itself.

This form of test has a particular advantage. The results of the Dickey–Fuller tests presented in Table 1 for the individual countries indicated that it was impossible to reject the hypothesis of a unit root in almost all cases. If, as argued, this failure is due to the few degrees of freedom available in these series, then pooling the data for all or most of the countries becomes a necessity. Given the number of countries involved, however, multivariate tests like that of Johansen, even though more powerful, are a good deal less tractable than the alternative panel-data versions of conventional univariate unit root tests. The tests that I conducted were based on the following fixed effects model:

$$\langle 9 \rangle \qquad \Delta(q_{it} - \bar{q}_{i.}) = \lambda(q_{it-1} - \bar{q}_{i.}) + \sum_{j=1}^{J} \beta_j \Delta(q_{it-j} - \bar{q}_{i.}) + u_{it},$$

where the bar and dot again indicate a country mean, Δ indicates a first difference, and the lag length J varies from 1 to 4 years. As usual, the value of λ is the focal point. An estimated value of λ that is significantly different than zero results in a rejection of the null hypothesis of a unit root.[12]

Table 5 contains the results of these tests. As is readily apparent from the table, the unit-root null is consistently rejected, something that was not possible in the much less powerful tests for the 22 countries viewed individually. The estimates of λ in the four test regressions are all negative and range in value from -0.46 to -0.29. These in turn imply estimated half-lives of the deviations of real exchange rates from their long-term equilibria ranging from slightly over 1 year to slightly over 2 years. The latter in particular is close to the estimates obtained in other studies and not far removed from the estimates described above, which were derived from comparisons of regressions run with varying period-average growth rates.

III. Conclusions

As the 1980s drew to a close, purchasing power parity and exchange-rate models that relied on it stood in near total discredit. Evidence since then has been a good deal more favorable to PPP, but the bulk of that evidence has come from examination of long historical data sets. Whether the conclusions of such studies are applicable to the current float continues to be debated.

This study suggests that they are. The stylized facts of the floating rate period considered most inimical to PPP were the high volatility of real exchange rates and strong positive correlation between real and nominal exchange rates observed in month-to-month, quarter-to-quarter and even year-to-year data. The cross-country results reported in this paper, however, point to an important additional fact, that both the volatility of real exchange rates and their correlation with nominal rates diminish greatly—the latter eventually disappearing—as data frequency is reduced.[13] Other tests show, moreover, that real exchange rates are better characterized as mean-reverting than as following unit-root processes. On a long-term average basis, therefore, purchasing power parity turns out to have retained a good deal of empirical usefulness.

While this finding stands in contrast to prevailing beliefs about exchange-rate behavior during the current float, it is, nevertheless, very similar to what researchers have uncovered in investigations of long historical time-series data. One inference to be drawn from these results is that the alleged differences between the behavior of real exchange rates under the current float and under earlier regimes are in fact more apparent than real. A second implication has to do with the theory of exchange rate determination. Theories that focus exclusively on permanent real shocks as the major force driving exchange rates under the float clearly miss much of what actually transpired. Transitory influences, nominal and perhaps also real, appear to have played a much more important role than commonly believed.

What those influences are, and in particular what caused the substantial real dollar appreciation in the first half of the 1980s and the equally dramatic depreciation thereafter, are questions that need to be answered. A related question has to do with what Rogoff (1996) has recently termed 'the purchasing power parity puzzle'—why the adjustment of real exchange rates to equilibrium takes so long.

Notes

1. Shiller and Perron (1985) and Hakkio and Rush (1991) make this argument in the contexts of unit-root tests and Engle–Granger tests of cointegration, respectively.

Frankel (1986), Lothian (1990), and Lothian and Taylor (1997) among others, contain discussions of the problem in the specific context of exchange-rate behavior.

2. Recent studies using long-term time series that reach this conclusion include Abuaf and Jorion (1990), Lothian (1990), and Diebold *et al.* (1991) and Lothian and Taylor (1996). Frankel (1986) and Edison (1987) are somewhat earlier studies of this type.
3. Stockman (1990), for example, views the slow estimated speeds of adjustment as evidence inconsistent with monetary overshooting.
4. See Grilli and Kaminsky (1991). Based on the differences in variance ratios computed for the float and earlier regimes, they conclude that the float was in fact different and conjecture that the cause of the difference was the greater incidence and severity of real shocks.
5. Lothian and Taylor (1996) pursue this approach and find a stable relationship for the float vis-à-vis earlier historical experience, 1791 to 1973 in the case of dollar–sterling and 1805–1973 in the case of franc–sterling.
6. Abuaf and Jorion (1990) is the first such application with which I am familiar. They report results that are consistent with mean reversion of real exchange rates, but generally too weak to reject the unit-root hypothesis with any high degree of confidence for the countries that they examine under the float. Several recent studies, which I have come across while this paper was in draft form, have been more successful in rejecting the unit-root hypothesis with panel data. These include Frankel and Rose (1996), Jorion and Sweeney (1996), and Oh (1996). Mark (1995) reaches similar conclusions using somewhat different econometrics.
7. A notable exception is Frankel and Rose (1996).
8. The exchange rates are yearly averages as listed in either line rf or line rh of the *International Financial Statistics*; the figures for the cost-of-living indexes are yearly averages as listed in line 64 of that publication.
9. Papell (1996) has recently shown that in panel data for real exchange rates such as those used here, consistent rejection of the unit-root null requires a relatively large group of countries. The dominance of the large upward and then downward movement in the average dollar real exchange rate plotted in Figure 1 suggests why this is the case.
10. The similarity between the two suggests that one potential problem with PPP regressions of this sort–bias due to the errors in variables problem plays no role in these data. If $\hat{p}-\hat{p}^*$ alone were measured with error, the slope coefficient in the regression of $\hat{p}-\hat{p}^*$ on, $\hat{e}$ call it β_e, would be an unbiased and consistent estimate of the true coefficient. Alternatively (and less realistically) if only $\hat{e}$ were measured with error, then, $1/\beta_p$, the reciprocal of the slope coefficient in the regression of $\hat{e}$ on $\hat{p}-\hat{p}^*$, would provide an unbiased and consistent estimate of the true coefficient. If both variables were measured with error, then these two estimates, β_e and $1/\beta_p$, would provide lower and upper bounds on the true coefficient. Table 2 reports values for β_e and β_p of 1.034 and 0.957, respectively, which in turn imply lower and upper bounds of 1.03 and 1.05, which leaves little room for any such bias.

    A second potential problem is simultaneity, but this also seems to be of little consequence in actuality. Assuming that average rates of money supplygrowth over the period are exogenous, we can use the average differential in money supply growth in place of the average inflation differential as the independent variable in the nominal exchange rate regression to get around this problem. Doing so resulted in an almost identical coefficient to that for the inflation differential (1.06 vs 1.03) suggesting that any bias on this score is also minimal.
11. The initial sample for the 6-year averages was 1973–79 and for the 9-year averages 1973–1982. To construct additional samples I proceeded sequentially, dropping the initial year and adding the next (previously omitted) year in each instance—dropping 1973 and adding 1980, for example, in the case of the second of the 6-year samples; dropping 1973 and adding 1983 in the case of the second of the 9-year samples.
12. Frankel and Rose (1996), Jorion and Sweeney (1996), and Oh (1996), conduct similar tests but with different data sets than mine. Since one of the panel data sets used by

Oh contains an almost identical number of country-times-year observations—23 countries over the years 1973–1989 in Oh's sample vs 22 countries over the years 1974 to 1990 in mine—I am able to make use of the critical values that he derived via Monte Carlo simulations.

13. It is important to point out that these results in no way appear to be simply an artifact of this particular data sample. As shown above, they continue to hold when different temporal and spatial subsamples are analysed. They actually improve somewhat when the period is extended forward (see Lothian and Simaan, 1996). Finally, they are consistent with the results reported in other recent studies of the float by researchers using largely differing econometric techniques and data sets to mine (see the references cited in note 6).

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