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A CENTURY OF PURCHASING-POWER PARITY

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A Century of Purchasing-Power Parity
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ABSTRACT

This paper investigates purchasing-power parity (PPP) since the late nineteenth century. I collected data for a group of twenty countries over one hundred years, a larger historical panel of annual data than has ever been studied before. The evidence for long-run PPP is favorable using recent multivariate and univariate tests of higher power. Residual variance analysis shows that episodes of floating exchange rates have generally been associated with larger deviations from PPP, as expected; this result is *not* attributable to significantly greater persistence (longer halflives) of deviations in such regimes, but is due to the larger shocks to the real-exchange rate process in such episodes. In the course of the twentieth century there was relatively little change in the capacity of international market integration to smooth out real exchange rate shocks. Instead, changes in the size of shocks depended on the political economy of monetary and exchange-rate regime choice under the constraints imposed by the trilemma.

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Introduction

What new findings can this paper claim to offer given the wealth of research on PPP in the past? It first should be noted that empirical support for PPP has waxed and waned over the years. From an historical standpoint, there have been numerous studies of PPP for various countries over the period in question, some covering a particular era or monetary regime. McCloskey and Zecher (1984) argued that PPP worked very well under the Anglo-American gold standard before 1914. Diebold, Husted, and Rush (1991) explored a very long run of nineteenth century data for six countries, and found support for PPP based on the low-frequency information lacking in short-sample studies. Abauf and Jorion (1990) studied a century of dollar-franc-sterling exchange rate data and verified PPP; Lothian and M. P. Taylor (1996) found the same for *two* centuries of dollar-franc-sterling data. Lothian (1990) also found evidence that real exchange rates were stationary for Japan, the U.S., the U.K., and France for the period 1875-1986, although yen exchange rates exhibited only trend-stationarity—an oft-repeated finding that the real yen exchange rate has appreciated over the long run against *all* currencies. In full length monographs, both Lee (1978) and Officer (1982) found strong evidence in favor of PPP based on analysis of long time-series running from the pre-1914 gold standard to the managed float of the 1970s.[1]

Of late, new studies have appeared in abundance. In their recent comprehensive review of the purchasing-power parity literature, Froot and Rogoff (1995) could declare that what was a "fairly dull research topic" only a decade ago has recently been the focus of substantial controversy and the subject of a growing body of literature. Recent empirical research, mostly based on the time-series analysis of short spans of data for the floating-rate (post-Bretton Woods) era led many to conclude that PPP failed to hold, and that the real exchange rate followed a random walk, with no mean-reversion property. However, a newly emerging literature exploits more data and higher-powered techniques, and claims that, in the long run, PPP does indeed hold: it appears from these studies that real exchange rates exhibit mean reversion with a half-life of deviations of four to five years (M. P. Taylor 1995; Froot and Rogoff 1995). The newer findings use various steps to expand the size of samples used to test PPP. As noted, it has been possible to use much longer-run time series for certain individual countries, spanning a century or more; typically such exercises have concentrated on more-developed countries with good historical data availability (for example, U.S., Britain, France). Alternatively, researchers have expanded the data for the recent float or postwar periods cross-sectionally to exploit the additional information in panel data (Wei and Parsley 1995; Frankel and Rose 1995; Pedroni 1995; Higgins and Zakrajšek 1999).

It is still too early to say whether the revisionist PPP findings will prove robust, and already challenges to this interpretation have emerged. One may find fault with the ways in which cross-section information and panel methodologies have been applied (O'Connell 1996, 1998). Some have noted that the inferences based on panel methods are sensitive to sample selection, and many results appear sensitive to the choice of base country, for example, the U.S. versus Germany (Papell 1995; Wei and Parsley 1995;

[1] Obviously, this paper builds on a very strong foundation of historical work by a number of scholars, covering various countries in different time periods. Other studies of long run data are numerous (Frankel 1986; Edison 1987; Johnson 1990; Glen 1992; Kim 1990).

Edison, Gagnon, and Melick 1995). Others caution that detecting a unit root in time series may be complicated by the fact that price indices can be viewed as the sum of a stationary tradable relative-price component and a non-stationary non-tradable relative-price component (Engel 2000; Ng and Perron 1999). This finding echoes the venerable Balassa-Samuelson objection to the pure PPP hypothesis based on differential rates of productivity growth in traded and non-traded goods sectors (Balassa 1964; Samuelson 1964). Of course, such long-run trends may be purely deterministic (Obstfeld 1993).

The distinction of the present study is to bring very recent empirical innovations to a longer span of historical data, both to investigate the robustness of the recent findings and to explore the historical evolution of PPP. I proceed as follows. Section 2 introduces the real exchange rate data and preliminary analysis shows that old-style univariate tests cannot reject the unit root null. Section 3 examines a multivariate test of PPP by M. P. Taylor and Sarno (1998). In the search of a more powerful test, Section 4 applies the univariate efficient tests of Elliott, Rothenberg, and Stock (1996). We find that long-run PPP can be supported in all cases with allowance for deterministic trends.

The importance of the long-run trends is explained in Section 5 where I model the dynamics of real exchange rates at different times in the twentieth century. Four regimes are investigated, the gold standard 1870–1914, the interwar period 1914–45, the Bretton-Woods era 1946–71, and the recent float 1971–96. The important quantitative differences found are in residual variance, and the floating regimes exhibit much larger shocks to the real exchange rate process, accounting for the much larger deviations from PPP during these eras. Thus, the history of PPP in the twentieth century shows, surprisingly, that there was relatively little change in the ability of international market integration to smooth out real exchange rate shocks. Instead, I argue, the changes in the variance of the shocks reveal a great deal about the differing degrees to which monetary policy was kept in check or not by commitment mechanisms (under fixed rates) or their absence (under floating). In light of this, I end with a discussion that relates these findings to the question of the political economy of monetary and exchange-rate regime choice under the constraints imposed by the macroeconomic policy trilemma.

Data and Preliminary Analysis

The data consist of annual exchange rates E_{it}, measured as domestic currency units per U.S. dollar, and price indices P_{it}, measured as consumer price deflators—or, when they are not available, GDP deflators. We will refer to the log levels of these variables, denoted $e_{it} = \log E_{it}$ and $p_{it} = \log P_{it}$. The index $i = 1, \ldots, 20$ covers the set of countries Argentina, Australia, Belgium, Brazil, Canada, Chile, Denmark, Finland, France, Germany, Greece, Italy, Japan, Mexico, Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, the United Kingdom, and the United States. The index t runs over the set of years from 1850 to 1996, but a complete cross-section of 20 exchange rates does not exist before 1892, the starting date of the Swiss series.[2]

[2] In constructing the dataset I have relied on standard sources. After 1948 the series are taken from the IMF's *International Financial Statistics* on CD-ROM. The principal pre-1948 price sources are the statistical volumes of Brian Mitchell. For the provision of electronically-compiled price and exchange rate data from these and other sources I am grateful to Michael Bordo. An appendix containing the data and full documentation is available from the author upon request.

Given these data, some preliminary transformations and tests were performed. Let the U.S. dollar-denominated price level of country i at time t be denoted by $R_{it} = P_{it}/E_{it}$, with $r_{it} = \log R_{it} = p_{it} - e_{it}$. As an initial step, missing data were filled in for each series. In all cases, this amounted to imputing a value to a few wartime years for certain countries, using linear interpolation on r_{it}. This yields a balanced 20×105 panel of data from 1892 to 1996.

Such an interpolation procedure may be *ad hoc*, but it was deemed necessary to give any stationarity test a fair chance on this data, since, in several cases, the missing data appear after explosive inflations during which real exchange rate often depreciated. Without interpolation in these periods, any subsequent reversion back toward the mean (or trend) in this variable would be missed by any estimation procedure, and a bias against stationarity would result. An important example would be the wide divergence in real exchange rates in the 1930s following the collapse of the gold standard; this episode was followed by war, leading to many missing observations in the data, and thus much of the reversion of these divergent real exchange rates toward PPP during and after the war would be omitted from the sample absent any interpolation.

With interpolations complete, the real exchange rate series was generated two ways: first, relative to the U.S. dollar, as $q_{it} = r_{it} - r_{\mathrm{US},t}$; and second, relative to the "world" ($N = 20$) basket of currencies, as $q_{it}^W = r_{it} - r_{it}^W$, where $r_{it}^W = \frac{1}{N-1}\sum_{j \neq i} r_{jt}$.[3] The second definition follows O'Connell (1996), and may help us avoid problems associated with the choice of the United States as a base country.[4]

The complete series q_{it} and q_{it}^W for all 20 countries are shown in Figure 1. One way to test the PPP hypothesis is to ask: are these real exchange rates stationary, that is, mean-reverting? A cursory inspection suggests that for many countries real exchange rates have been fairly stable over the long run, and we might expect to easily support the hypothesis of stationarity. Nonetheless, our eyes are drawn to certain cases where there appears to be a long-run trend or random walk. Here, the most obvious and well-known problem would be the case of Japan, but similar symptoms of drift or nonstationarity might also be perceived for Switzerland, Brazil, and in some other countries' experience in specific periods, such as interwar Germany and Italy. Clearly, a powerful statistical test will be needed to resolve this question.

We can begin analysis using more traditional unit root tests. Table 1 shows the results of applying the augmented Dickey-Fuller (ADF) unit root tests to the univariate real exchange rate series, and the results are expected given the findings in the previous literature.[5] In many cases, the unit root null cannot be rejected. Even allowing a trend to be present does not seem to help very much, and the null is not rejected in most cases. However, a simple OLS regression on a constant and a trend seems to indicate that, for at least some of the series, a deterministic trend might be present; this trend component is a sizeable 1.5% per annum in the case of Japan, 0.74% per annum for Switzerland,

[3] Ideally, one might prefer to use trade-weighted real exchange rates, but such data do not exist in the form of annual time series for the entire twentieth century for a wide sample of countries. Future research would need to be directed to original sources to collate the necessary bilateral trade volumes, and this would be a significant undertaking.

[4] This discussion was omitted in O'Connell (1998).

[5] Lag lengths in the ADF tests were chosen by the Lagrange Multiplier (LM) criterion for residual serial correlation, allowing up to a maximum of 6 lags.

Figure 1: A Century of Real Exchange Rates

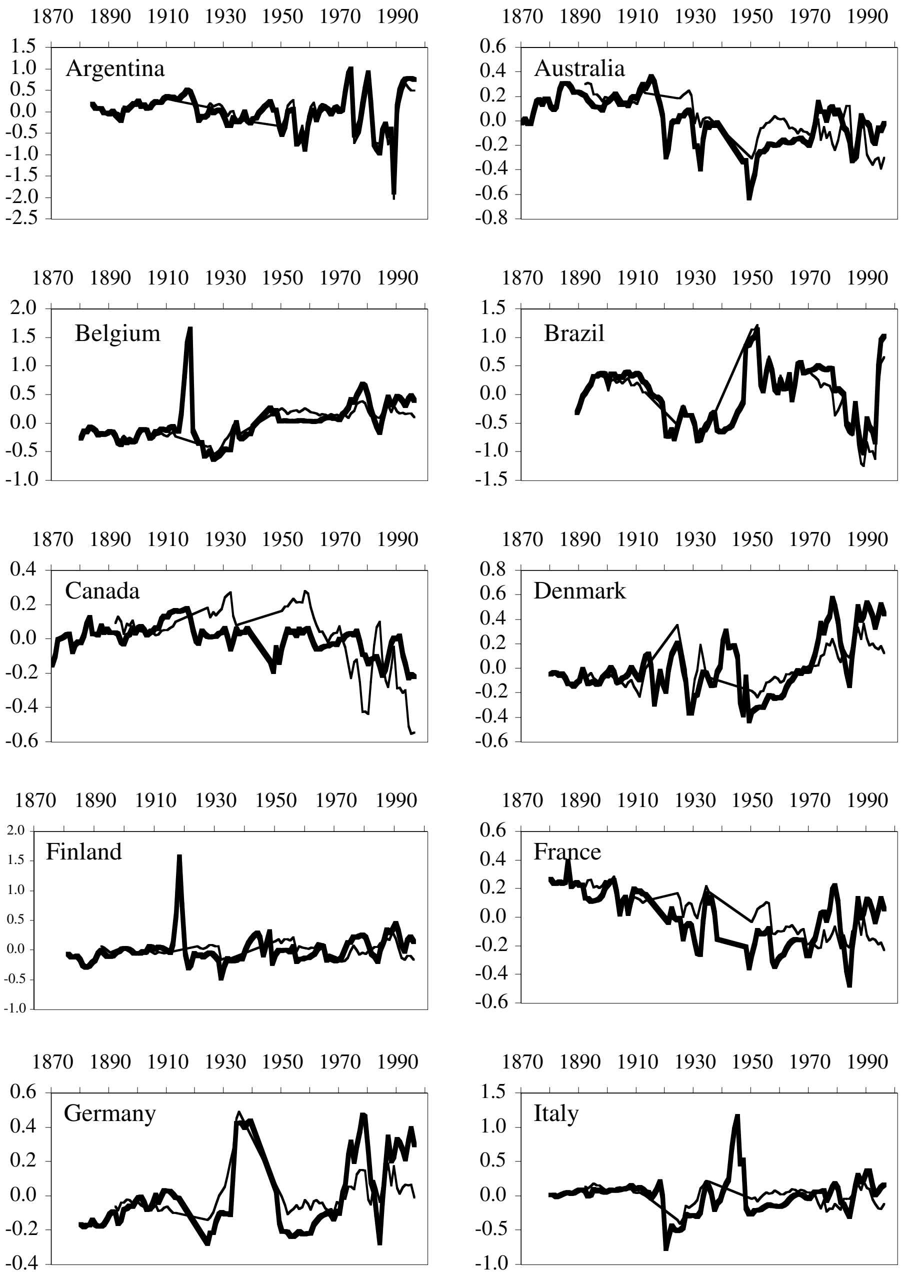
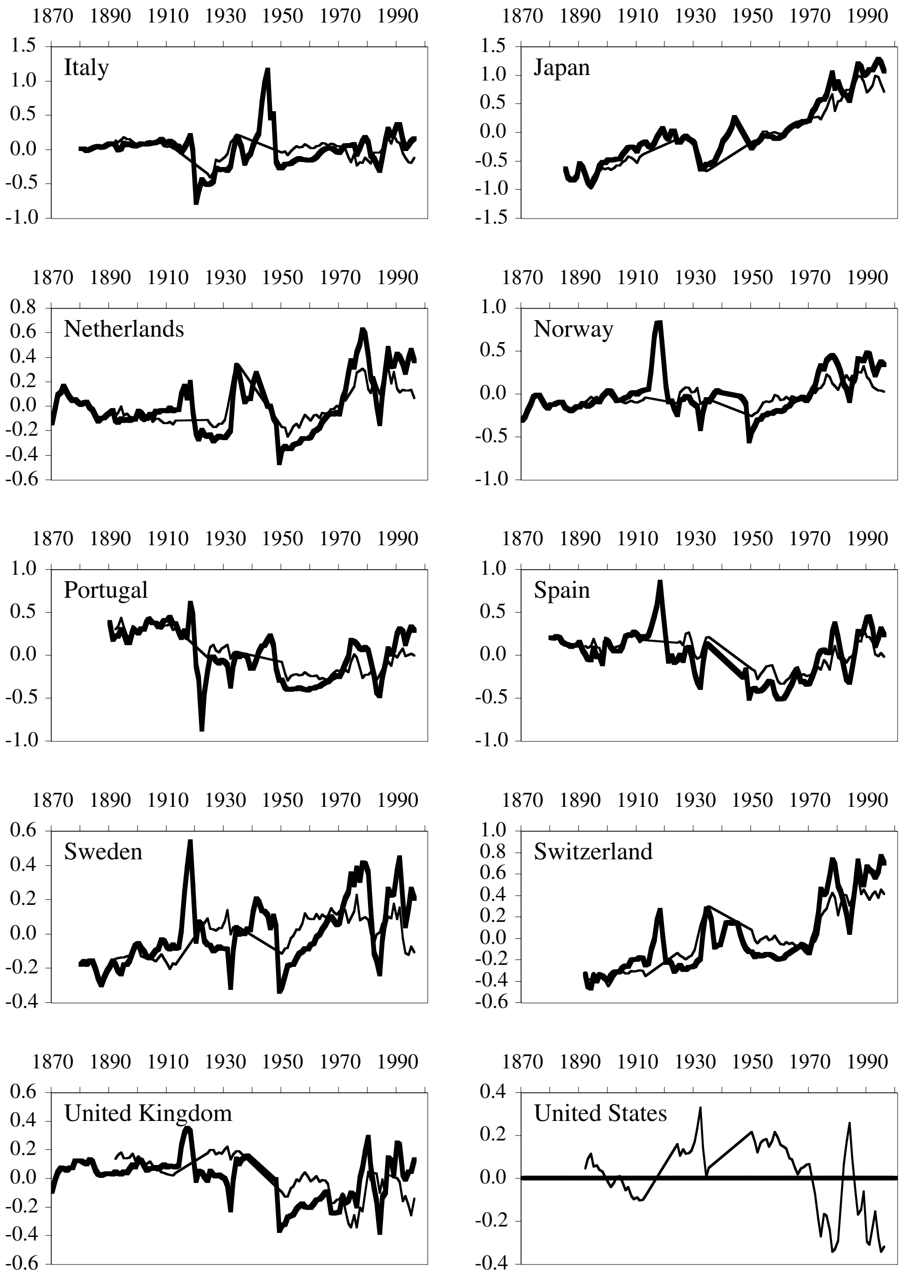


Figure 1: A Century of Real Exchange Rates (continued)



Notes and Sources: See text and appendix. The thicker line shows q_{it}, the real exchange rate relative to the U.S. dollar. The thinner line shows q_{it}^{W}, the real exchange rate relative to the "world" ($N = 20$) basket of currencies.

Table 1: Preliminary Data Analysis

	T	Demeaned ADF	Demeaned LM	Detrended ADF	Detrended LM	OLS Trend
Base: United States						
Argentina	113	-4.85 ***	0	-4.81 ***	0	-0.0013
Australia	127	-2.44	0	-3.04	0	-0.0030 ***
Belgium	117	-3.25 **	0	-3.85 **	0	0.0053 ***
Brazil	108	-2.74 *	1	-2.70	1	0.0000
Canada	127	-2.59 *	0	-3.76 **	0	-0.0011 ***
Denmark	117	-2.16	0	-2.77	0	0.0033 ***
Finland	116	-4.50 ***	0	-4.63 ***	0	0.0016 **
France	117	-3.54 ***	1	-4.18 ***	1	-0.0030 ***
Germany	117	-2.96 **	1	-3.28 *	1	0.0029 ***
Italy	117	-3.27 **	0	-3.27 *	0	0.0005
Japan	112	-0.29	0	-1.99	0	0.0151 ***
Mexico	111	-2.96 **	0	-3.91 **	0	-0.0065 ***
Netherlands	127	-2.00	0	-2.27	0	0.0022 ***
Norway	127	-2.41	0	-2.61	0	0.0025 ***
Portugal	107	-2.60 *	0	-2.46	0	-0.0037 ***
Spain	117	-2.34	0	-2.25	0	-0.0019 ***
Sweden	117	-2.78 *	0	-3.44 **	0	0.0032 ***
Switzerland	105	-1.93	1	-3.43 **	1	0.0083 ***
United Kingdom	127	-3.19 **	0	-3.41 *	0	-0.0014 ***
Base: "World" Basket						
Argentina	105	-5.19 ***	0	-5.32 ***	0	-0.0029 **
Australia	105	-2.38	0	-3.56 **	0	-0.0045 ***
Belgium	105	-3.44 **	0	-4.07 ***	0	0.0051 ***
Brazil	105	-2.32	0	-2.26	0	-0.0017
Canada	105	-1.24	0	-2.04	0	-0.0033 ***
Denmark	105	-2.84 *	0	-3.35 *	0	0.0028 ***
Finland	105	-4.70 ***	0	-4.69 ***	0	-0.0002
France	105	-2.59 *	0	-3.96 **	0	-0.0038 ***
Germany	105	-1.80	0	-1.80	0	0.0017 ***
Italy	105	-3.20 **	0	-3.19 *	0	-0.0004
Japan	105	-0.97	0	-2.35	0	0.0150 ***
Mexico	105	-1.45	2	-4.44 ***	1	-0.0084 ***
Netherlands	105	-2.13	0	-2.54	0	0.0027 ***
Norway	105	-2.50	0	-2.51	0	0.0012 **
Portugal	105	-2.93 **	0	-3.71 **	0	-0.0051 ***
Spain	105	-2.03	0	-2.30	0	-0.0030 ***
Sweden	105	-2.67 *	0	-2.54	0	0.0017 ***
Switzerland	105	-1.05	0	-2.75	0	0.0074 ***
United Kingdom	105	-2.00	0	-2.88	0	-0.0032 ***
United States	105	-2.34	0	-2.53	0	-0.0013 **

Notes and Sources: See text and appendix. T is the sample size. ADF is the augmented Dickey-Fuller statistic with LM the lag length selected by the Lagrange Multiplier criterion. Demeaned is the case where each series is replaced by the residuals from a regression on a constant. Detrended is the case where the regression is on a constant and a linear trend. Trend is the OLS estimate of the linear trend. Finite-sample critical values are shown based on 4,000 simulations of the null; $*$ denotes significance at the 10% level; $**$ denotes significance at the 5% level; $***$ denotes significance at the 1% level.

but is small (no more than half a percent per year) in all other cases. All in all, we are left with the conclusion that although some of the series may be $I(1)$, many are $I(0)$ and most, in addition, have some deterministic drift. This impression is given whether one uses the U.S as a base country, or one measures real exchange rates relative to the "world" basket.

One traditional response to such findings has always been to fault these tests for their lack of power, and to point to the fact that, with slow convergence speeds, the autoregressive parameter might be very close to unity, and one would need a very long span of data to reject the null (Frankel 1990). With over a century of data, we might just have sufficient span to have a reasonably powerful test, but we are still unable to find broad evidence of stationarity. The recent literature suggests two possible directions: the use of multivariate or panel methods, and the use of more efficient univariate tests. We pursue both routes, to see which, if any, might lend support to the PPP hypothesis.

A Multivariate Test

If PPP holds among a set of $N+1$ countries, this would imply that every single log real exchange rate, for all $N(N+1)/2$ bilateral pairs, would be stationary. One axiomatic property of well-defined PPP measures is *base country invariance*. That is, the concept of PPP must be invariant to the choice of base country. Thus we may, without loss of generality, take a particular choice of a base country, and then consider the N bilateral log rates relative to that country.[6]

An elegant test for the stationarity of these, and hence all, log rates was proposed by M. P. Taylor and Sarno (1998). They note that a necessary and sufficient condition that all N series be stationary would be the existence of N independent cointegrating vectors among the series (Engle and Granger 1987). Conversely, if one takes the null to be the absence of such a condition, that is, the existence of *any* nonstationary real exchange rates, the null would correspond to a situation where the N series had fewer than N cointegrating vectors.

These hypotheses permit some simple tests based on the cointegration methods of Johansen (1988, 1991). To briefly review this approach, let $\mathbf{q}_t = (q_{1t}, \ldots, q_{Nt})$ be the $N \times 1$ vector of real exchange rates at time t. Under the PPP hypothesis, all N components of $\mathbf{q}$ are $I(0)$. The error-correction representation for the dynamics of $\mathbf{q}$ is

$$\Delta\mathbf{q}_t = \mathbf{\Gamma}_1\Delta\mathbf{q}_{t-1} + \ldots \mathbf{\Gamma}_{k-1}\Delta\mathbf{q}_{t-k+1} + \mathbf{\Gamma}_k\mathbf{q}_{t-k} + \boldsymbol{\mu} + \boldsymbol{\omega}_t.$$

The $N \times N$ matrix $\mathbf{\Gamma}_k$ has rank equal to the number of cointegrating vectors, so the null hypothesis of *one or more* nonstationary series is: $H_0 : \text{rank}(\mathbf{\Gamma}_k) < N$, and the alternative hypothesis where *all* series are stationary is: $H_1 : \text{rank}(\mathbf{\Gamma}_k) = N$. Since full rank of $\mathbf{\Gamma}_k$ would imply that all of the eigenvalues are nonzero, a test of H_1 against a null of H_0 amounts to a test of the restriction that the smallest eigenvalue λ_N of the estimated $\mathbf{\Gamma}_k$ matrix is zero. The Johansen likelihood ratio test statistic for this case is $JLR = -T\ln(1-\lambda_N)$ where JLR has an asymptotic distribution that is $\chi^2(1)$.

[6]All log real exchange rates between bilateral pairs can then be derived, assuming arbitrage amongst all cross-rates in the exchange market, as linear combinations of the set of N log rates for the given base country.

Table 2: Johansen Likelihood Ratio Test

Base: World	*SIM*	Demeaned *JLR*	*p*	power	Detrended *JLR*	*p*	power
Group 1: Europe							
Belgium, France, Germany, Italy, U.K. Netherlands	4,032	1.77	[.37]	0.25	5.04	[.01]	0.02
excluding Netherlands	3,968	2.99	[.18]	0.31	4.77	[.02]	0.02
excluding Netherlands and Germany	3,840	3.80	[.13]	0.43	8.09	[.00]	0.02
Group 2: Scandinavia							
Denmark, Finland, Sweden, Norway	3,840	3.11	[.22]	0.54	4.09	[.14]	0.29
excluding Norway	3,584	7.41	[.02]	0.60	7.25	[.03]	0.40
Group 3: Iberia and Latin America							
Argentina, Brazil, Mexico, Portugal, Spain	3,968	4.06	[.08]	0.33	4.86	[.01]	0.02
Group 4: Other							
Australia, Canada, Switzerland, Japan	3,840	0.65	[.59]	0.12	4.34	[.03]	0.01
excluding Switzerland and Japan	3,072	6.53	[.03]	0.23	7.64	[.01]	0.07

Notes and Sources: See text and appendix. *JLR* is the Johansen Likelihood Ratio test statistic. The finite-sample significance level is based on *SIM* simulations estimated under the assumptions of the null. The power of the 5% test is based on 4, 096 simulations estimated under the assumptions of the alternative.

The chief merit of this test may be the clean specification of null and alternative hypotheses. Other mutlivariate tests based on panel methods, such as the multivariate form of the ADF test, may reject a unit root null when only some of the series are stationary, but not all N.[7] A multivariate approach also places further restrictions on any empirical framework. Given the base-country-invariance postulate, the structure imposes k lags of the $\Delta\mathbf{q}_t$ in all equations of the system. Such a restriction is commonly not a feature of univariate tests of PPP using single-country exchange rate series, yet it ought to be present if we are really thinking in terms of a joint hypothesis test involving the stationarity of all the series taken together. But what lag choice should one make? In the previous section, the tests performed in Table 1 report a variety of lag lengths selected by the Lagrange Multiplier criterion. By inspection, we note that, the maximal lag length is two, so I chose $k = 2$ lags for the *JLR* test, as the minimal lag length that should eliminate serial correlation from all univariate series.

The results of applying the Taylor-Sarno *JLR* test are shown in Table 2, with finite-sample significance levels and power calculations. The results shown are those for the real exchange rate relative to a “world” base, but the results using the U.S. as a base country are similar and are omitted to save space. The width of the panel is potentially a problem here: we have $N = 20$ where M. P. Taylor and Sarno had $N = 4$. Empirical implementation would be inefficient, clumsy, and costly if I attempted to estimate a 20-equation VAR for the dynamic equation, so I elected to work with four subsets with between four and six countries in each.[8]

[7] Other tests may suffer from a “missing middle” — the null is that all N series are $I(1)$, but the alternative of interest is that all N series are $I(0)$. This structure fails to recognize that there are many intermediate cases, where some series are stationary and some are not. This seems particularly relevant to our empirical task, since the data in Figure 1 and the preliminary data analysis in Table 1 suggest that we might well have such intermediate cases.

[8] I thank a referee for suggesting this partitioning approach. Finite-sample significance levels and power are derived by simulation. In simulations with $N \leq 6$ series there are $2^N - 1$ combinations of $I(1)$ and $I(0)$ series that satisfy the null, and only one, with every series $I(0)$, that satisfies the alternative. I ran $2^{12} = 4,096$ simulations on each test, with each null combination simulated 2^{12-N} times.

The *JLR* tests are somewhat favorable to the PPP hypothesis for most countries, but only when allowance is made for a trend. In the cases with trend, stationarity is accepted at a better than 10% significance level, except for the case of Norway which just breaks that significance threshold. As expected, the power of test is low and it falls dramatically when a trend is included. This reflects a common problem in the literature, namely the low power of tests to detect trend stationarity in favor of a unit root null. In the tests without a trend, stationarity is rarely accepted, except for Denmark, Finland, Sweden, Australia and Canada, just 5 countries out of 20. The results suggest that the *JLR* tests, in this particular sample, suffer from the weak power problem identified by M. P. Taylor and Sarno. Accordingly, we might shift attention to a univariate approach that uses the most efficient tests possible and is flexible enough to handle slowly-evolving deterministic components.

A Univariate Test

The most powerful univariate unit root tests available at present are the generalized-least-squares (GLS) versions of the Dickey-Fuller (DF) test due to Elliott, Rothenberg, and Stock (1996). The tests are of broad applicability since they apply to cases where the series have: (*i*) no trend; (*ii*) a deterministic constant term $d_t = (1)$; and (*iii*) a deterministic constant term and drift $d_t = (1, t)$. We are, as always, working with index numbers in PPP tests, and we also might want to allow for possible deterministic trends in the spirit of Balassa-Samuelson, so the DF-GLS test is very relevant.

In the DF-GLS test, the series z_t to be tested is replaced in the ADF regression by $\tilde{z}_t = z_t - \hat{\beta}' d_t$, where $\hat{\beta}'$ is a GLS estimate of the coefficients on the deterministic trends d_t. That the DF-GLS test dominates others is shown via a local-to-unity asymptotic approach, and the power envelope is close to the frontier. The unit-root PPP controversy hangs on being able to pin down an autoregressive parameter ρ that is less than, but often very close to, unity. Hence, the DF-GLS test is an ideal tool for PPP testing.[9]

Table 3 shows the results of applying the DF-GLS tests to our real exchange rate data. The format repeats that of Table 1. Four cases are considered: using the U.S. and the "world" basket as a base; and using the series demeaned and detrended. These results offer powerful support for the PPP hypothesis in the twentieth century. In all cases without detrending the null of a unit root is rejected, and in most cases even with a trend, though the test is less powerful there. Hence, with some allowance for the possibility of slowly-evolving long-run trends, I conclude that PPP has held in the long run over the twentieth century for my sample of 20 countries.[10]

If PPP holds in the long run, it is no longer productive to devote further attention to the stationarity question. The more important and interesting problem is to explain what drives the short-run dynamics of real exchange rates.[11] That is, how do we account for the amplitude and persistence of deviations from PPP, in different time periods and in different countries in the last century?

[9] The DF-GLS test gives support for PPP in the post-Bretton Woods era (Cheung and Lai 1998).

[10] It would be desirable to follow up this study in the future with tests based on higher-frequency data. Still, that we can find evidence in favor of PPP with annual series is very encouraging indeed, given the biases introduced by temporal averaging in historical data (Taylor 2001).

[11] The same conclusion was reached by Higgins and Zakrajšek (1999).

Table 3: DF-GLS Tests

	Base: United States		*Base: "World" Basket*	
	Demeaned	Detrended	Demeaned	Detrended
Argentina	-4.79 ***	-4.76 ***	-5.13 ***	-5.31 ***
Australia	-2.45 **	-3.10 **	-2.47 ***	-3.59 ***
Belgium	-3.23 ***	-3.89 ***	-3.45 ***	-4.10 ***
Brazil	-2.70 ***	-2.79 **	-2.34 **	-2.36
Canada	-2.60 **	-3.98 ***	-1.47 *	-2.29
Denmark	-2.20 **	-2.86 **	-2.85 ***	-3.39 ***
Finland	-4.49 ***	-4.67 ***	-4.67 ***	-4.72 ***
France	-3.54 ***	-4.15 ***	-2.62 ***	-4.00 ***
Germany	-2.94 ***	-3.30 ***	-1.81 *	-1.83
Italy	-3.28 ***	-3.30 ***	-3.20 ***	-3.22 **
Japan	-0.93 **	-2.12	-1.55 ***	-2.37 *
Mexico	-2.96 ***	-3.95 ***	-1.62 *	-4.46 ***
Netherlands	-2.06 **	-2.33	-2.15 **	-2.58 *
Norway	-2.46 ***	-2.65 *	-2.52 **	-2.54 *
Portugal	-2.62 ***	-2.63 **	-2.94 ***	-3.75 ***
Spain	-2.35 **	-2.35 *	-2.05 **	-2.34
Sweden	-2.79 ***	-3.47 ***	-2.68 ***	-2.73 **
Switzerland	-2.13 **	-3.41 ***	-1.39 *	-2.79 **
United Kingdom	-3.18 ***	-3.46 ***	-2.06 **	-2.92 **
United States	—	—	-2.37 **	-2.61 *

Notes and Sources: See Table 1, text, and appendix. The lag length is selected by the Lagrange Multiplier criterion. Demeaned is the case where each series is replaced by the residuals from a regression on a constant. Detrended is the case where the regression is on a constant and a linear trend. Finite-sample critical values are shown based on 4,000 simulations of the null. $*$ denotes significance at the 10% level; $**$ denotes significance at the 5% level; $***$ denotes significance at the 1% level. The critical values corresponding to these significance levels are $(-1.62, -1.95, -2.58)$ for the demeaned series and $(-2.57, -2.89, -3.48)$ for the detrended series, respectively. See Elliott, Rothenberg, and Stock (1996).

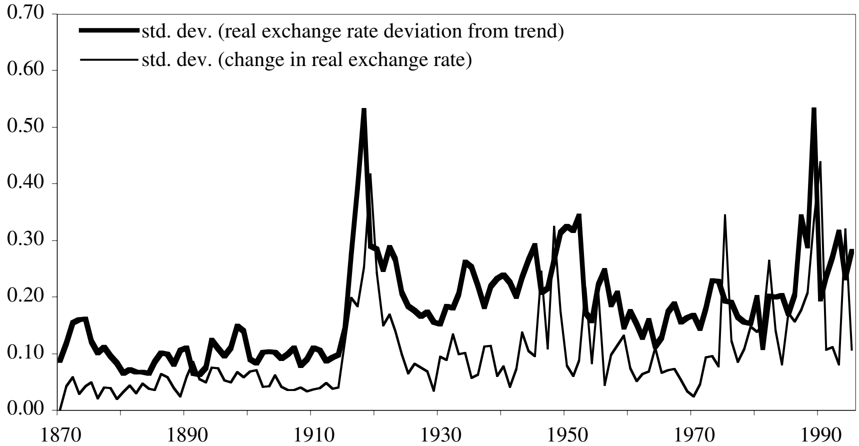
An Overview of PPP in the Twentieth Century

In this section, given the earlier findings, deviations from PPP will be measured relative to the equilibrium real exchange rate. As we have seen, it is necessary to allow for slowly-evolving deterministic trends. As an empirical matter, they are usually found to be "small." However, their omission would undoubtedly upset any study of the deviations of real exchange rates over the very long run.[12] Accordingly, I will, for the remainder of this paper, consider the dynamics of detrended real exchange rates in an attempt to measure the reversion speed toward equilibrium.

The first question to ask is: what have been the extent of deviations from PPP over the long run? One way to answer this question is to examine volatility via the size of changes in the real exchange rate Δq_{it}, since, according a mean-reversion theory, this change would be proportional to the deviation from equilibrium plus some error. Another approach would be to detrend the series q_{it} and examine the deviations of the resulting detrended level q_{it}, that is, the error-correction term. For a cross section of countries, the extent of these deviations at a given time t can be measured by the standard deviations $\sigma(\Delta q_{it})$ and $\sigma(q_{it})$. Figure 2 shows these measures for our entire sample and both exhibit similar trends.

[12] A trend of, say, 0.5% per annum might make little difference over a one to ten year horizon, but over one hundred years, if such a correction were left out, then log deviations from equilibrium could be mismeasured by an additive shift of 0.5, or in levels by a multiplicative shift of 65%.

Figure 2: Real Exchange Rate Volatility and Deviations from Trend



Notes and Sources: See text and appendix.

Real exchange rate deviations and volatility were relatively small prior to 1914 under the classical gold standard regime, as expected. The interwar period was a major turning point; deviations became much larger as many exchange rates began to float or stay fixed for only a few years. There was some reduction in deviations after 1945, notably during the heyday of Bretton Woods during the 1960s. Once the floating rate era began in the 1970s, deviations and volatility once again rose. This chronology offers some *prima facie* reasons to view changes in the exchange rate regime as a major determinant of real exchange rate behavior, an idea we will keep in mind.

Although we can now see from the data where and when deviations have been large or small, we would like to know why they were large or small at particular times. In an autoregressive model, any changes in the properties of the deviations can only be attributed to two essential causes: either the dynamic process is subject to (stochastic) shocks of different amplitude; or else the process itself exhibits different patterns of (deterministic) persistence. To investigate this more fully, then, we need to apply and estimate a model. Given that we are taking trend stationarity as given, based on earlier findings, Table 4 reports the results of fitting an error-correction model to the detrended U.S.-based real exchange rate q_{it}, with a specification

$$\Delta q_{it} = \beta_0 q_{it} + \beta_1 \Delta q_{i,t-1} + \beta_2 \Delta q_{i,t-2} + \epsilon_t .$$

The coefficients β_1 and β_2 are not reported; columns labeled *i* and *t* indicate the samples, including pooled samples(P) across both countries and time periods; periods correspond to the exchange rate regimes, Gold Standard (G), Interwar (I), Bretton Woods (B), and Float (F); halflives in years are reported (H); and significance levels are reported for tests of pooling across periods (p1) and countries (p2).[13]

[13]The lag choice $k = 2$ was sufficient based on LM tests in all cases except the cross-country pooled samples. A uniform lag structure was imposed to facilitate pooling tests.

Table 4: A Model of Real Exchange Rates

i	*t*	β_0	s.e.	R^2	*T*	*H*	p1	p2	*i*	*t*	β_0	s.e.	R^2	*T*	*H*	p1
P	P	-0.21	(0.01)	.11	2,293	3.4	.00	.00	ITA	P	-0.25	(0.06)	.15	114	3.6	.00
P	G	-0.21	(0.03)	.13	633	3.1		.01	ITA	G	-0.36	(0.15)	.28	31	1.9	
P	I	-0.24	(0.03)	.20	640	3.1		.88	ITA	I	0.00	(0.14)	.09	32	-21.8	
P	B	-0.43	(0.03)	.28	520	1.5		.63	ITA	B	-0.54	(0.06)	.81	26	1.0	
P	F	-0.41	(0.04)	.19	500	2.1		.99	ITA	F	-0.32	(0.16)	.38	25	2.3	
ARG	P	-0.47	(0.10)	.20	110	1.5	.96		JPN	P	-0.09	(0.04)	.15	109	9.3	.07
ARG	G	-0.09	(0.12)	.09	27	6.0			JPN	G	-0.27	(0.14)	.26	26	2.9	
ARG	I	-0.18	(0.10)	.14	32	4.1			JPN	I	-0.08	(0.06)	.29	32	8.9	
ARG	B	-0.47	(0.20)	.22	26	1.4			JPN	B	-0.25	(0.11)	.69	26	1.6	
ARG	F	-0.59	(0.24)	.24	25	1.2			JPN	F	-0.35	(0.15)	.23	25	1.8	
AUS	P	-0.18	(0.05)	.11	124	4.7	.15		MEX	P	-0.25	(0.07)	.15	108	2.4	.47
AUS	G	-0.19	(0.07)	.30	41	5.2			MEX	G	-0.09	(0.14)	.31	25	3.6	
AUS	I	-0.34	(0.13)	.20	32	2.3			MEX	I	-0.15	(0.09)	.16	32	6.2	
AUS	B	-0.23	(0.13)	.14	26	4.0			MEX	B	-0.45	(0.17)	.25	26	1.6	
AUS	F	-0.53	(0.18)	.32	25	1.6			MEX	F	-0.45	(0.23)	.27	25	1.1	
BEL	P	-0.30	(0.07)	.21	114	2.6	1.00		NLD	P	-0.11	(0.04)	.13	124	7.8	.09
BEL	G	-0.31	(0.14)	.16	31	2.3			NLD	G	-0.12	(0.06)	.13	41	6.8	
BEL	I	-0.31	(0.14)	.21	32	2.5			NLD	I	-0.23	(0.11)	.21	32	3.8	
BEL	B	-0.29	(0.10)	.39	26	2.1			NLD	B	-0.21	(0.13)	.13	26	3.2	
BEL	F	-0.34	(0.13)	.39	25	3.0			NLD	F	-0.37	(0.14)	.37	25	2.1	
BRA	P	-0.13	(0.06)	.07	105	4.8	.37		NOR	P	-0.15	(0.04)	.20	124	6.2	.08
BRA	G	-0.46	(0.14)	.38	22	2.0			NOR	G	-0.31	(0.09)	.39	41	2.7	
BRA	I	-0.27	(0.10)	.24	32	2.4			NOR	I	-0.20	(0.09)	.36	32	4.8	
BRA	B	-0.48	(0.18)	.28	26	1.2			NOR	B	-0.35	(0.16)	.18	26	2.0	
BRA	F	-0.22	(0.17)	.10	25	3.9			NOR	F	-0.42	(0.14)	.34	25	2.0	
CAN	P	-0.20	(0.06)	.10	124	3.4	.20		PRT	P	-0.17	(0.06)	.10	104	5.2	.05
CAN	G	-0.10	(0.10)	.05	41	3.2			PRT	G	-0.13	(0.14)	.11	21	3.0	
CAN	I	-0.19	(0.11)	.22	32	3.3			PRT	I	-0.48	(0.16)	.25	32	1.7	
CAN	B	-0.25	(0.15)	.19	26	2.2			PRT	B	-0.18	(0.07)	.50	26	3.3	
CAN	F	-0.35	(0.13)	.33	25	5.0			PRT	F	-0.17	(0.11)	.32	25	5.2	
DNK	P	-0.15	(0.05)	.10	114	4.9	.00		SPA	P	-0.13	(0.04)	.15	114	7.1	.01
DNK	G	-0.60	(0.19)	.38	31	1.5			SPA	G	-0.21	(0.14)	.10	31	3.0	
DNK	I	-0.36	(0.14)	.28	32	2.2			SPA	I	-0.27	(0.11)	.30	32	3.5	
DNK	B	-0.55	(0.18)	.35	26	0.8			SPA	B	-0.41	(0.15)	.29	26	1.3	
DNK	F	-0.38	(0.14)	.35	25	2.4			SPA	F	-0.22	(0.10)	.48	25	2.8	
FIN	P	-0.39	(0.08)	.28	113	1.8	.21		SWE	P	-0.23	(0.06)	.19	114	3.3	.82
FIN	G	-0.21	(0.11)	.25	30	4.3			SWE	G	-0.30	(0.12)	.31	31	2.9	
FIN	I	-0.40	(0.16)	.35	32	1.8			SWE	I	-0.28	(0.14)	.21	32	2.4	
FIN	B	-0.57	(0.22)	.51	26	0.5			SWE	B	-0.27	(0.15)	.21	26	2.3	
FIN	F	-0.41	(0.14)	.38	25	2.0			SWE	F	-0.37	(0.15)	.29	25	2.6	
FRA	P	-0.22	(0.06)	.17	114	3.3	.03		SWI	P	-0.13	(0.05)	.21	102	5.0	.04
FRA	G	-0.51	(0.23)	.27	31	0.9			SWI	G	-0.38	(0.24)	.45	19	0.7	
FRA	I	-0.44	(0.15)	.33	32	1.7			SWI	I	-0.29	(0.12)	.37	32	3.1	
FRA	B	-0.64	(0.20)	.34	26	1.3			SWI	B	-0.28	(0.06)	.60	26	2.1	
FRA	F	-0.36	(0.14)	.35	25	2.4			SWI	F	-0.36	(0.14)	.34	25	1.7	
GER	P	-0.10	(0.04)	.23	114	6.8	.13		UKG	P	-0.20	(0.06)	.10	124	3.6	.10
GER	G	-0.19	(0.12)	.16	31	3.5			UKG	G	-0.22	(0.10)	.14	41	1.9	
GER	I	-0.06	(0.05)	.31	32	16.0			UKG	I	-0.27	(0.14)	.21	32	2.6	
GER	B	-0.23	(0.06)	.56	26	2.3			UKG	B	-0.42	(0.13)	.35	26	1.5	
GER	F	-0.36	(0.15)	.33	25	2.2			UKG	F	-0.42	(0.19)	.20	25	1.7	

Notes and Sources: See text and appendix. The country abbreviations are: ARG Argentina; AUS Australia; BEL Belgium; BRA Brazil; CAN Canada; DNK Denmark; FIN Finland; FRA France; GER Germany; ITA Italy; JPN Japan; MEX Mexico; NLD Netherlands; NOR Norway; PRT Portugal; SPA Spain; SWE Sweden; SWI Switzerland; UKG United Kingdom. Samples are P Pooled; G Gold Standard; I Interwar; B Bretton Woods; F Float.

Table 5: Model Halflives and Error Disturbances

	Halflife					*SEE*				
	P	G	I	B	F	P	G	I	B	F
Pooled	3.4	3.0	3.1	1.6	2.1	.14	.05	.15	.11	.20
Argentina	1.8	7.2	4.0	1.6	1.5	.33	.08	.12	.25	.64
Australia	4.3	3.0	2.6	4.1	2.1	.08	.04	.11	.08	.08
Belgium	2.6	2.5	2.5	2.6	3.1	.19	.07	.35	.04	.11
Brazil	4.9	0.8	2.6	1.5	3.3	.26	.07	.15	.30	.39
Canada	3.9	6.0	2.8	2.7	3.7	.04	.04	.04	.05	.04
Denmark	4.4	1.5	2.4	0.8	2.8	.10	.04	.11	.11	.11
Finland	1.9	3.9	2.0	0.6	2.5	.16	.04	.26	.12	.10
France	3.2	1.0	2.0	1.6	2.7	.08	.06	.08	.06	.10
Germany	7.2	2.7	11.7	4.5	2.5	.07	.03	.08	.04	.11
Italy	3.6	1.5	—	2.1	2.5	.14	.03	.20	.09	.10
Japan	8.4	3.2	8.8	3.9	2.2	.09	.07	.09	.04	.12
Mexico	2.1	6.2	5.3	2.2	1.3	.17	.10	.15	.13	.27
Netherlands	6.4	6.3	3.5	3.9	2.6	.08	.03	.10	.08	.11
Norway	5.3	3.4	4.2	2.4	2.7	.09	.03	.13	.09	.09
Portugal	4.7	4.2	2.2	4.1	4.2	.13	.06	.19	.05	.10
Spain	5.8	3.0	3.3	2.1	3.2	.11	.07	.13	.09	.10
Sweden	3.0	2.9	2.4	2.1	2.8	.09	.03	.11	.08	.12
Switzerland	5.2	0.7	3.0	1.8	2.1	.09	.03	.10	.03	.12
United Kingdom	3.7	3.1	2.5	2.3	2.1	.08	.02	.08	.07	.13
Mean	4.3	3.3	3.7	2.4	2.6	.13	.05	.14	.10	.16
Standard Deviation	1.8	1.9	2.5	1.1	0.7	.07	.02	.07	.07	.14
Median	4.1	3.0	2.8	2.1	2.6	.10	.04	.12	.08	.11

Notes and Sources: See text and appendix. Samples are P Pooled; G Gold Standard; I Interwar; B Bretton Woods; F Float.

Note that these results are often for very short spans of data, so that we are not using the coefficient β_0 as a basis for a stationarity test. Rather, we now have a maintained hypothesis of long run trend stationarity based on the earlier tests. The pooling restrictions are not always rejected, but sufficiently often that is seems safest to treat this as a heterogeneous panel, and examine the nature of the dynamics in different periods and countries. This is pursued in Table 5, by focusing on the two key features—one random, one not—that generate PPP deviations: the halflife of disturbances, calculated from the estimated model via (deterministic) forecast; and the variance of the (stochastic) error disturbances $SEE = \sigma_\epsilon$.[14]

The striking aspect of these results are the relatively small variations in halflives across the four exchange-rate regimes. There are notable exceptions. One is Italy in the interwar period, where the estimated root is explosive on this restricted sample; also, interwar Germany has slow reversion which may not be surprising given the aftermath of hyperinflation in the 1920s and the extensive controls on the economy in the 1930s (see Figure 1).[15] Still, all the other halflives in the table are in the low single digits as

[14]For a simple motivation of this rough division of sources of deviations, consider an AR(1) process for the real exchange rate, $q_t = \rho q_{t-1} + \epsilon_t$. The unconditional variance of q_t is $\mathrm{Var}(q) = \sigma_\epsilon^2/(1-\rho^2)$. The halflife is a simple function if the autoregressive parameter, $H = \ln 0.5/\ln\rho$. Thus, the numerator of $\mathrm{Var}(q)$ is a function of the size of the (stochastic) shocks, and the denominator a function of the (deterministic) halflife. With higher order processes the separation is not so clean, but the intuition is the same.

[15]Tests for PPP in the 1930s for Britain, U.S., France and Germany were undertaken by Broadberry and M. P. Taylor (1988). Consistent with the present interpretation, they found PPP except for bilateral exchange rates involving the mark, a result attributed to the extensive controls in the German economy.

measured in years. The mean and median halflives hover around two to three years, a timeframe even more favorable to rapid PPP adjustment than most recent empirical studies. The variation in halflives around the mean or median is small, around one or two years in most cases. There is evidence of only a modest decline in halflives after World War Two, with a drop from 3.5 years to 2.5 on average, a decline of about one third. In sum, we have found a new, quite provocative, and remarkable result. Looking across the twentieth century, and despite considerable differences in institutional arrangements and market integration across time and across countries, the *deterministic* aspects of persistence of PPP deviations have been fairly uniform in the international economy.[16]

What, then, accounts for the dramatic changes in deviations from PPP during the twentieth century seen in Figure 2? As one might guess, it is the *stochastic* components that have to do most of the work to account for this given the fairly flat halflife measures. Under the gold standard we find $SEE = .05$ on average, that is, a 5% standard deviation for the stochastic shocks. This rises by a factor of three to 14% on average in the interwar, then falls by a third to 10% under Bretton Woods, before climbing by over one-half to 16% under the float. Of course, there are some notable outliers here, such as the Latin American economies that experience hyperinflation in the postwar period. We should also note that, due to lack of accurate, synchronized data, the German hyperinflation of the early 1920s is omitted from the data in this study, and is covered by interpolation.

To reinforce the point, Figure 3 shows a scatterplot of $\sigma(\Delta q_t)$ versus σ_ϵ for the AR model fitted to each country during each regime. Given the model is linear, we may write $\sigma(\Delta q_t) = f(\rho_j)\sigma_\epsilon$ where the ratio f is a function of all the AR coefficients ρ_j. With no persistence $f = 1$, and more persistence causes an increase in $f > 1$. What is noteworthy here is how f has been uniform and almost constant over the twentieth century. The correlation of $\sigma(\Delta q_t)$ and σ_ϵ has been 0.99 across all regimes. From the regression we see that a forecast of $\sigma(\Delta q_t)$ assuming $f = 1.1$ yields an R^2 of 0.9 and a tiny standard error of 0.01. As we surmised, the persistence of the processes has played little role here little, and changes in the stochastic shocks explain virtually all changes in the volatility in the real exchange rate across space and time.

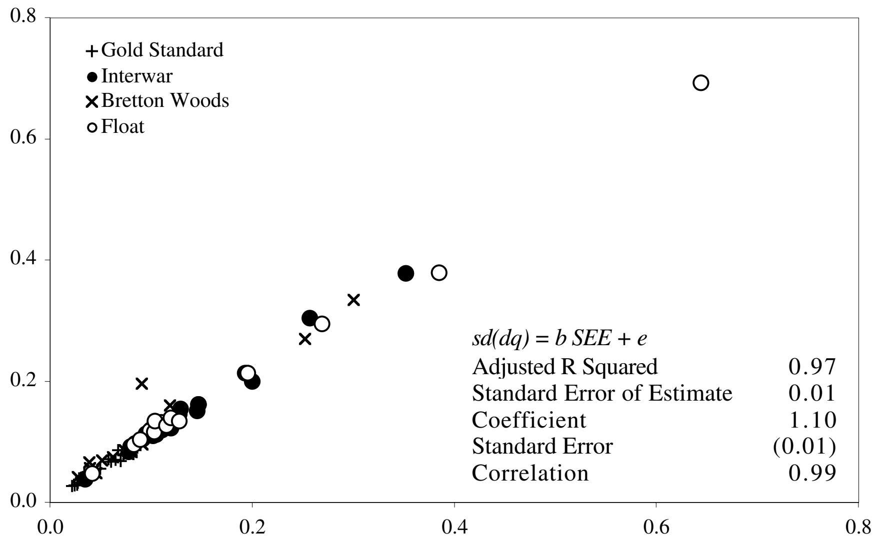
The error disturbances tell a consistent story, revealing much larger shocks to the real exchange rate process under floating-rate regimes than under fixed-rate regimes. This result has been observed in contemporary data, but this study is the most comprehensive long-run analysis, based on more than a century of data for a broad sample of countries. Of particular historical note is the emergence of the interwar period as an important turning point, an era when PPP deviations shifted to dramatically higher levels.[17] Given the vast changes in institutions and market structure over a hundred years or more, the relationship of real exchange rate deviations to the monetary regime now looks like a robust stylized *historical* fact.

A final piece of evidence reinforces this notion. One approach to explaining real exchange rate deviations in cross section has been to try to disengage the effects of

[16]Another study that examines reversion to PPP across different monetary regimes is Parsley and Popper (1999). They focus only on postwar data for the period 1961–92 in 82 countries, but this encompasses a wide range of exchange rate arrangements. They find only slightly faster reversion under the dollar peg, about 12% per year, versus pure floating, at 10% per year. This is consistent with the findings in this paper.

[17]On the interwar period as a turning point see Obstfeld and A. M. Taylor (1998; 2001). Interwar studies of PPP find results consistent with these findings (Eichengreen 1988; M. P. Taylor and McMahon (1988).

Figure 3: Real Exchange Rate Volatility ($\sigma(\Delta q_t)$) versus SEE (σ_ϵ)



Notes and Sources: See text and appendix. Vertical axis: $\sigma(\Delta q_t)$. Horizontal axis: σ_ϵ.

geography and currencies. Engel and Rogers (1996, 1998, 1999) have shown that although "border effects" do matter, a very large share of deviations from parity across countries is accounted for by the effect of currencies, that is, by nominal exchange rate volatility.[18] We can follow a similar tack here, by looking at the sample variances for our four regimes and for each of the twenty countries. Of course, unlike Engel and Rogers we cannot make within country comparisons, but we do have a somewhat more controlled experiment: using an historical sample, as opposed to the post-Bretton Woods era, we do obtain much greater sample variation in exchange rate volatility, even as "geography"—needless to say—has remained constant.

Table 6 tabulates real and nominal exchange rate volatility in the various subsamples. Under the gold standard we see low real and nominal volatility among those countries that clung hard to the rules of the game (those with zero nominal volatility); but for other countries, as the nominal volatility rose, so did the real volatility (examine, for example, Japan and Switzerland, then Mexico, Portugal and Spain, and finally Brazil and Argentina). Overall the cross country correlation is 0.74. Under the mostly-floating interwar period a similar story can be told, although many more real shocks were present in the form of terms-of-trade disturbances and financial crises, so it is perhaps not surprising to see the correlation fall to 0.52. Another reason that the correlations might be somewhat less than one in the early twentieth century is that price

[18] An example of their approach would be to regress $\sigma(\Delta q_{it})$ on $\sigma(\Delta e_{it})$ and measures of distance plus a "border" dummy (equal to one when the locations are in different countries). Within Europe, for the 1980s and 1990s, they find there is an almost one-to-one pass through from $\sigma(\Delta e_{it})$ to $\sigma(\Delta q_{it})$ (the coefficient is 0.92), and an inspection of the summary statistics for each is sufficient to convey the message (Engel and Rogers 1999, Tables 2 and 3A).

Table 6: Real Versus Nominal Exchange Rate Volatility

	G		I		B		F	
	$\sigma(\Delta q)$	$\sigma(\Delta e)$	$\sigma(\Delta q)$	$\sigma(\Delta e)$	$\sigma(\Delta q)$	$\sigma(\Delta e)$	$\sigma(\Delta q)$	$\sigma(\Delta e)$
Pooled	6	5	18	16	16	18	22	47
Argentina	8	13	12	10	27	25	69	112
Australia	4	1	13	10	8	8	10	10
Belgium	7	4	43	15	5	4	13	13
Brazil	9	15	16	15	33	39	38	101
Canada	4	2	4	4	5	4	5	4
Denmark	5	2	12	13	13	12	12	12
Finland	5	0	30	21	16	18	12	12
France	7	0	11	21	8	8	12	13
Germany	3	0	9	9	5	6	13	13
Italy	3	2	20	29	20	20	12	13
Japan	9	5	11	9	4	3	13	13
Mexico	11	7	15	8	14	13	29	35
Netherlands	4	2	11	10	8	8	13	12
Norway	4	1	17	15	10	8	10	10
Portugal	7	7	21	27	7	3	12	14
Spain	7	7	18	15	9	10	13	14
Sweden	3	0	12	10	9	8	13	13
Switzerland	5	4	11	10	4	2	14	14
United Kingdom	3	0	9	9	9	8	13	14
Corr($\sigma(\Delta q)$,$\sigma(\Delta e)$)								
by regime	0.74		0.52		0.99		0.94	
all regimes	0.87							

Notes and Sources: See text and appendix.

flexibility was almost certainly higher in this earlier epoch, a result noted in international studies of business-cycle fluctuations.[19] In the postwar period the correlation is very strong, 0.99 under Bretton Woods and 0.94 under the float for our sample. In the float, Brazil and Argentina pose problems for the correlation because of their hyperinflation experiences—episodes when, again, large price adjustments went in tandem with nominal exchange rate movements. The correlation for the twenty countries over all regimes is 0.87, and the message I take from these results is that the dominant source of PPP failure is nominal exchange rate volatility, that is, the nature of the monetary regime.[20]

Finally, we might ask, why was this pattern of real and nominal exchange rate volatility observed in twentieth century history, and what implications should this have for our research? The empirical measures shown here appear very consistent with historical changes in monetary regimes, the associated record of institutional changes, and the tools from the political-economy nexus that have been invoked to explain them. The widely accepted account of these major regime shifts relies on what Obstfeld and

[19]See the survey of these issues in Basu and A. M. Taylor (1999).

[20]Do all international relative prices move up and down together as per the aggregate real exchange rate movement, or do they show different patterns that are less well correlated with nominal exchange rate volatility? Absent detailed disaggregated data, we cannot show, like Engel and Rogers did (1995, 1996), how much of these PPP deviations are common to all goods' relative prices as a source of deviations from the law of one price (LOOP). This would be an excellent topic for future research. However, unless contradicted by an array of large and offsetting LOOP deviations for various goods that virtually cancel out—an unlikely outcome—the patterns thus far are entirely consistent with the view that deviations from LOOP are similarly traceable to deviations from aggregate PPP, which, in turn, are in large part determined by the nature of monetary shocks, rather than barriers to trade or geography.

A. M. Taylor (1998, 2001) term the *macroeconomic policy trilemma*. This trilemma is the well-known conflict facing policymakers when choosing between three competing objectives, (*i*) a fixed exchange rate, (*ii*) capital mobility, and (*iii*) activist monetary policy, where only two out of three are feasible. Under this schema, the gold standard saw countries forsake monetary policy (*iii*). The interwar was a period when either controls, sacrificing (*ii*), or devaluations, sacrificing (*i*), were employed. Bretton Woods was a system of limited capital mobility, entailing the loss of (*ii*). The float brought back capital mobility at the expense of fixed rates sacrificing (*i*). Several measures of capital mobility in the twentieth century accord with the trilemma view of history, and historical accounts paint a similar picture once we examine institutional change and the actions of policymakers more closely (Eichengreen 1996).

The tight relationship between monetary volatility and real exchange rate volatility sustains doubts about meaningful macroeconomic models that impose short-run money neutrality. In the long run PPP holds, and so money appears to be neutral at that horizon; but the fact that short-run PPP deviations may be large, and seem very closely associated with monetary shocks, suggests a role for nominal rigidities. Since the real exchange rate is a combination of price levels and exchange rates, another way to restate the conclusion is that inflation volatility and nominal exchange rate volatility—each one a monetary phenomenon in itself—are jointly nonneutral in the sense that they are correlated with a real effect, the size deviations from PPP (as noted by Cheung and Lai 2000). The above correlations would then be consistent with a view that nominal exchange rates can adjust very quickly even as other prices in the economy move more sluggishly, an assumption common to many international macroeconomic models of older and newer vintages (Dornbusch 1976; Obstfeld and Rogoff 1996). Further study will be needed to incorporate these dynamics into an econometric PPP model and measure them in historical (and contemporary) samples, but it does seem that monetary time series would be extremely important as an explanatory variable, despite their considerable endogeneity problems. In short we leave this study with the strong suspicion that for the most part, to coin a phrase, deviations from PPP are always and everywhere a monetary phenomenon.

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