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## Assessing convergence to purchasing power parity: a panel study for ten OECD countries

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*Suggested Citation:* Meier, Carsten-Patrick (1997) : Assessing convergence to purchasing power parity: a panel study for ten OECD countries, Weltwirtschaftliches Archiv, ISSN 0043-2636, Mohr, Tübingen, Vol. 133, Iss. 2, pp. 297-312

This Version is available at:

<https://hdl.handle.net/10419/1719>

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# Assessing Convergence to Purchasing Power Parity: A Panel Study for Ten OECD Countries

By

Carsten-Patrick Meier

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## I. Introduction

**T**he theory of purchasing power parity (PPP) holds that international arbitrage causes exchange rates and relative price levels of two countries to be strictly proportional. As this phenomenon has seldom been observed empirically, economists have started to ask whether prices and exchange rates do at least converge to strict proportionality in the long run by testing the corresponding real exchange rates for nonstationarity. By use of either very long-run time series of single real exchange rates or panel approaches that combine data from different real exchange rates, evidence has been found that long-run PPP does in fact hold. However, convergence to PPP seems to be “painfully slow” (MacDonald 1995: 481). Consensus estimates put the average rate at which convergence takes place at only 15 per cent per year, implying that the typical deviation from PPP has a half-life of about 4 years.<sup>1</sup>

This paper re-examines the evidence on long-run PPP for ten OECD countries by means of panel methods. It argues that for most OECD countries, convergence to PPP is indeed faster than previously indicated. The approach differs from existing studies in at least three ways. *First*, in contrast to panel methods used in other studies to test for convergence to PPP, the present approach does not rely on condensing the information contained in the various real exchange rate time series into a single average estimate. Instead, it allows to assess the rate at which the real exchange rate adjusts to a shock for each individual parity separately. This may provide useful additional infor-

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*Remark:* Financial support from the Thyssen Foundation is gratefully acknowledged.

<sup>1</sup> Surveys of the literature are given by Froot (1996), Froot and Rogoff (1995) and MacDonald (1995). An average half-life of 4 years is also found in a recent panel study by Frankel and Rose (1996).

mation on the adjustment processes in different countries that would otherwise be covered up in an average estimate. It may also indicate how much existing average estimates from panel studies are influenced by certain individual real exchange rates in the sample. *Second*, contrasting the work on the subject which uses aggregate price level-based real exchange rates, this study focuses exclusively on real exchange rates based on tradables prices. The rationale is that PPP can only be expected to hold for those goods and services which can be arbitrated internationally and is, therefore, to be studied on the basis of prices of these items only. Although long-run PPP has been shown to hold for aggregate price indexes, too, the estimated speed of adjustment towards parity will be biased as long as the relative price of nontradables to tradables is non-constant – as there are reasons to presume. *Third*, the analysis is confined to annual data since empirical as well as new theoretical considerations indicate that higher rates of convergence are likely to be found in low-frequency data. The analysis proceeds by first outlining the econometric technique, then the empirical results are presented and finally some conclusions are drawn.

## II. The Econometric Approach

Long-run purchasing power parity (PPP) is defined to hold if departures from strict PPP, as represented by deviations of the real exchange rate from its mean, vanish over time. Define the foreign country's real exchange rate  $q$  as  $q = e + p^* - p$  where  $e$  is the number of local currency units that buy one unit of foreign currency,  $p^*$  the foreign price level,  $p$  the local price level and all variables are expressed in logarithms. Since PPP requires the nominal exchange rate to change in proportion to the ratio of the foreign to local price levels, the real exchange rate represents departures from PPP. Long-run PPP requires these departures to disappear in the long run, that is, the time series of the real exchange rate is to return to its mean (of zero) in finite samples.

Consider the time series of the real exchange rate,  $q_t$ ,  $t = 1, \dots, T$ , to be generated by the autoregressive moving average process

$$\begin{aligned} q_t &= \varrho q_{t-1} + u_t - \theta u_{t-1}, \quad t = 1, \dots, T, \\ q_0 &= u_0, \end{aligned} \quad (1)$$

where  $\varrho$  and  $\theta$  are fixed parameters and the error term  $u_t$  is a stationary zero mean process. If  $\varrho = 1$ , the series follows a nonstationary random walk without drift, a process that does not return to its mean in finite

samples. Deviations from PPP would be considered as permanent. The usual strategy to determine empirically whether a time series follows a random walk, as proposed by Dickey and Fuller (1979, 1981), is to fit to the data an AR( $p$ )-model of the form

$$q_t = \mu + \rho q_{t-1} + \sum_{i=1}^p \Phi_i \Delta q_{t-i} + e_t \quad (2)$$

and to test if the autoregression coefficient  $\rho$  equals unity. A major drawback of this test is that it often does not reject the random walk hypothesis even if it is not true when  $\rho$  is close to 1 and the sample is small (Frankel 1986). Nonrejection of the random walk for real exchange rates, therefore, does not necessarily imply that deviations from PPP must be considered permanent when the sample contains only the 20 annual observations of the recent period of floating nominal exchange rates.

The statistical power of the Dickey-Fuller test can be increased by the use of larger samples. One approach is to use "panel"-based procedures by pooling time-series observations across various real exchange rates and evaluating the hypothesis that each individual series follows a random walk. In contrast to the alternative of extending the sampling period beyond the post-Bretton Woods period,<sup>2</sup> panel-based procedures allow the time-series information to remain confined to the period of floating nominal exchange rates. This seems desirable, because as the volatility of real exchange rates is higher under flexible than under fixed nominal exchange rates, mixing data from the two exchange rate regimes in PPP studies may bias the results towards acceptance of the PPP hypothesis.

Panel-based PPP tests have recently been applied by Frankel and Rose (1996), Wu (1996) and Abuaf and Jorion (1990). Frankel and Rose pool time-series and cross-section data for a large number of annual bilateral US real exchange rates based on consumer price indexes (CPI) for the post-war as well as for the post-Bretton Woods period, respectively. Applying the Dickey-Fuller test to the pooled data, they find that nonstationarity is rejected. For the post-1973 (to 1992) industrial country real exchange rates, the estimated autoregression coeffi-

<sup>2</sup> Typically, more than 80 annual observations are needed to reject the random walk for most of the important US dollar real exchange rates, so the sampling period has to be extended back to the beginning of the century. See Lothian and Taylor (1996) for a recent application. The problem with this approach is, of course, that results that are based to a large extent on data from a fixed exchange rates period may not hold for the present floating rates system.

cient implies a half-life of deviations from PPP about 4 years.<sup>3</sup> Abuaf and Jorion apply a somewhat different procedure to monthly (CPI-based) real exchange rates of the United States vis-à-vis ten OECD countries running from 1973 to 1987. They also reject nonstationarity. The estimated autoregression coefficients imply half-lives of a deviation from PPP between 3 and 5 years. Wu, in a similar study based on post-Bretton Woods US dollar real exchange rates of 19 OECD countries, also rejects the hypothesis that real exchange rates are not mean-reverting. He finds PPP deviations to have shorter half-lives: his estimates range from 2.7 years (for monthly CPI-based real exchange rates) to 1.9 years (for annual wholesale prices-based parities).

Evidently, in all three studies, increasing test power by resorting to panel procedures resulted in the finding that convergence to PPP is in fact present. However, this conclusion may have come at some cost: The procedures employed in these studies assume the autoregression coefficients  $\rho$  to be equal across the various real exchange rates. By averaging over the time-series information of a cross-section of real exchange rates valuable information on the behaviour of the individual real exchange rate is lost. In reality, the speed of adjustment of real exchange rates to exogenous shocks may well be different across countries, depending, among other things, on the shares of nontraded goods in the price indexes used, the degree of stickiness of the traded goods prices in the different countries, the countries' sizes, their openness to international trade as well as the various central banks' exchange rate policies. Likewise, different countries may simply have experienced different shocks to their real exchange rates over the last 20 years resulting in different adjustment experiences. If the adjustments of real exchange rates do indeed vary across countries, approaches which pool time-series and cross-section data to obtain a single autoregression estimate conceal valuable information about the adjustment processes in the different countries and are therefore inferior to an unrestricted estimation. Moreover, such an approach misspecifies the alternative hypothesis to the random walk making it more difficult to reject the latter because in fact a joint hypothesis is tested.

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<sup>3</sup> Frankel and Rose (1996) report a coefficient of 0.82 for their post-Bretton Woods/industrial countries sample which would imply that a shock to PPP is reduced by half after 3.5 years. However, they do not report adjustment for a small-sample bias. Simulations for the restricted GLS estimator used in the present paper showed that for a true  $\rho$  of 0.85 and a sample of 45 years of observations (the size of the cross-section did not significantly affect the bias) the bias is approximately 0.06. Adding this figure to the estimate of Frankel and Rose implies a half-life of a shock to PPP of 5.4 years.

Table 1 – *Correlations of Real Exchange Rates, 1973–1994*

	C	F	D	I	J	NL	N	CH	UK
C	1.00	<i>0.46</i>	<i>0.48</i>	<i>0.63</i>	<i>0.40</i>	<i>0.52</i>	<i>0.66</i>	<i>0.48</i>	<i>0.36</i>
F		1.00	0.98	0.93	<i>0.85</i>	0.98	0.93	0.94	<i>0.49</i>
D			1.00	0.90	0.90	0.99	0.91	0.94	<i>0.42</i>
I				1.00	<i>0.71</i>	0.91	0.96	0.90	<i>0.57</i>
J					1.00	<i>0.87</i>	<i>0.76</i>	<i>0.80</i>	<i>0.20</i>
NL						1.00	0.94	0.96	<i>0.48</i>
N							1.00	0.93	<i>0.60</i>
CH								1.00	<i>0.61</i>
UK									1.00

*Note:* Correlations refer to the logarithms of the real exchange rates vis-à-vis the US dollar based on value-added deflators for manufacturing. Values below 0.90 are printed in *italics*.

The least restrictive way to exploit the panel-type structure of real exchange rate data from different currency pairs is to estimate the Dickey-Fuller autoregressions as a system of seemingly unrelated regressions (SURE).<sup>4</sup> The procedure allows both the intercept and the autoregression coefficient to vary across the individual real exchange rate equations. As long as the disturbances of the equations are correlated to some extent, the Generalised Least Squares (GLS) estimator proposed by Zellner (1962), which takes these cross-equation correlations into account, will be more efficient than equation-by-equation OLS. The efficiency of the SURE procedure may be further enhanced by imposing cross-equation restrictions, requiring some or all of the coefficients to be the same across the various equations.<sup>5</sup> However, in the procedure applied below, this will only be the second step, as it should produce results comparable to the existing panel studies. New insights are primarily to be expected from the unrestricted estimation of the system.

Table 1 presents the matrix of correlations between a number of industrial countries' real exchange rates vis-à-vis the US dollar over the post-Bretton Woods period. The data is described in some detail below. Here, the important point to notice is the difference between the correlations. These differences can be expected to show up to some extent in the cross-equation correlations of the disturbances of the

<sup>4</sup> See Johnston (1984:396) for a comparison of the various procedures and models proposed for analyzing pooled time-series and cross-section data.

<sup>5</sup> This is the approach employed by Abuaf and Jorion (1990).

estimated equations. Since the efficiency of the GLS estimator relative to OLS increases with the cross-equation correlation of the disturbances, Table 1 may indicate for which real exchange rates the greatest gains in efficiency and test power may be expected from the GLS/SURE procedure. Evidently, these are the continental European countries for which the correlations amount to 0.9 and above. For the UK, Canada and Japan on the other hand only small gains in efficiency and power can be expected, if at all, as their real exchange rates do not correlate much with other countries' parities.

A problem with the GLS/SURE version of the Dickey-Fuller test is that the distribution of the test statistics is not known. One way to obtain the critical values for the test is by Monte Carlo simulation.<sup>6</sup> Artificial random-walk data is generated by a set of equations similar to (1), each with  $\mu = \theta = 0$  and  $\rho = 1$ . Since annual data for 9 real exchange rates from 1973 to 1994 are to be analyzed below, the data generation process consists of a system of 9 such equations, each generating 22 normally distributed observations with zero mean and exogenously given covariance matrix. The covariance matrix is taken to be the matrix of correlations between the real exchange rates in Table 1, as has been proposed by Abuaf and Jorion (1990). On the basis of this generated data, for which  $\rho$  is known to be unity, the multivariate version of (2) is then estimated by GLS/SURE. The estimate of  $\rho$  and the value of the  $t$ -statistic  $\tau = (\hat{\rho} - 1)/\hat{\sigma}_{\hat{\rho}}$ , where  $\hat{\sigma}_{\hat{\rho}}$  is the estimated standard error of  $\hat{\rho}$ , are noted. This experiment is replicated 5,000 times to obtain a sample distribution for  $\hat{\rho}$  and  $\tau$ .

The results of the Monte Carlo simulation are given in Table 2. Results of the unrestricted estimation are reported for the nine equations individually ordered according to which of the actual real exchange rates supplied the variance for the artificially generated data. For comparison, results for the single-equation Augmented Dickey-Fuller (ADF) test with 25 observations are presented, too. As argued above, the different cross-equation correlations of the disturbances result in different efficiency gains relative to OLS for the various real exchange rates. These differences are reflected in the simulation results concerning the critical values and the power of the test. Effectively, the sample is split in two groups with the continental European countries on the one hand and the remaining three countries on the other.

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<sup>6</sup> See Banerjee et al. (1993: 27ff) for a discussion and applications of Monte Carlo methods.

Table 2 – Monte Carlo Simulation for the GLS/SURE Test for Nonstationarity

Variance supplied by the US dollar real exchange of	Simulated critical values for the test statistic $\tau$ , significance level =		Probability (%) that the test statistic $\tau$ is smaller than the 5% critical value for a true value of $\rho =$		Bias assessment: estimated coefficient $\hat{\rho}$ for a true value of $\rho =$	
	5%	10%	0.70	1.00	0.70	1.00
Canada	-5.04	-4.47	39.8	4.6	0.53	0.77
France	-6.54	-5.34	72.5	5.1	0.53	0.76
Germany	-6.90	-6.31	72.6	4.8	0.53	0.77
Italy	-6.37	-5.89	74.3	5.3	0.53	0.77
Japan	-5.21	-4.76	36.1	4.7	0.52	0.76
Netherlands	-6.73	-6.17	53.3	4.6	0.53	0.76
Norway	-6.30	-5.69	65.5	5.7	0.52	0.77
Switzerland	-6.26	-5.70	70.6	5.2	0.53	0.77
UK	-5.34	-4.71	32.7	4.6	0.53	0.76
Restricted	-7.24	-6.74	53.3	4.9	0.57	0.82
ADF test	-3.00	-2.63	24.7	5.2	0.56	0.79

*Note:* All results are based on 5,000 replications and 22 observations per equation. The covariance matrix for the simulations is taken to be the matrix of correlations between the real exchange rates. Estimators and test statistics are obtained from GLS/SURE estimation of a system of regressions of the form  $q_{it} = \mu_i + \rho_i q_{it-1} + u_{it}$ ,  $i = 1, \dots, 9$ . The test statistic  $\tau = (\hat{\rho} - 1)/\hat{\sigma}_{\hat{\rho}}$  where  $\hat{\sigma}_{\hat{\rho}}$  is the estimated standard error of  $\hat{\rho}$ . Simulations assume  $\mu_i = 0$ . Critical values for the ADF test with 25 observations are taken from Fuller (1976: 373).  
 "Restricted": Estimation with the cross-equation restriction forcing all  $\hat{\rho}$  to be equal imposed.

The first two columns report critical values for  $\tau$  at conventional levels of significance. The critical values are considerably higher in absolute value than those for the single-equation Dickey-Fuller test. For the continental European countries they average, at the 5 per cent level, at about  $-6.5$  whereas for the UK, Canada and Japan the average is at about  $-5.2$ . When all nine equations are constrained to have the same autoregression coefficient, the critical value for rejecting nonstationarity at the 5 per cent level decreases to  $-7.24$ . The division between the continental European and the other three countries is also reflected in the statistical power of the GLS procedure. For a test with size 5 per cent and  $\rho = 0.7$ , a figure often found in autoregressions of real exchange rates at annual frequency, the probability that the null hypothesis of a random walk is rejected is about 70 per cent for the continental European countries. For the UK, Canada and Japan,



however, it is only about 35 per cent. Still, the GLS procedure is more powerful than the single-equation ADF test for which the comparable figure is 25 per cent. Thus, in any case, there are gains in efficiency to be expected from employing the GLS test. As regards the bias of the estimator in small samples, the table reports the mean value of the estimated  $\rho$  for true values of  $\rho = 1.0$  and  $\rho = 0.7$ . The simulations indicate that the GLS estimator is severely downward biased in a sample of only 22 observations. The bias is of the same magnitude as for the single-equation estimator and it is decreasing with decreasing  $\rho$ . For  $\rho = 0.7$  the average small sample bias is 0.173 for the unrestricted estimation. Imposing the cross-equation restriction reduces the bias to 0.134.

### III. Empirical Results

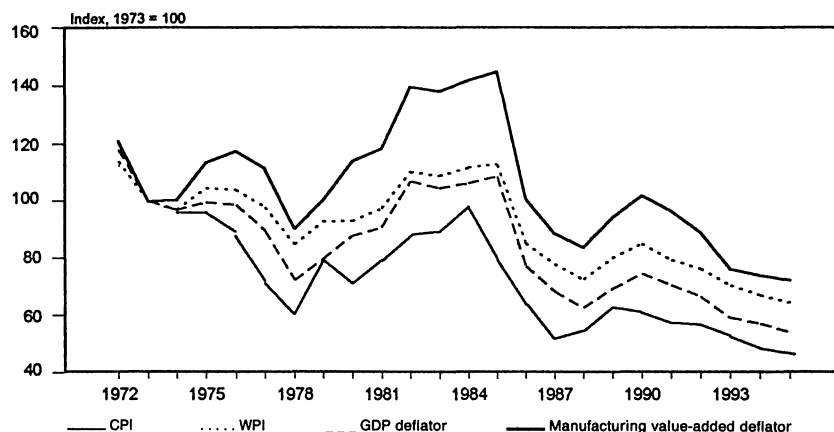
The empirical analysis focuses on the bilateral real exchange rates of the United States vis-à-vis Canada, France, Germany, Italy, Japan, the Netherlands, Norway and the United Kingdom for the post-Bretton Woods period. In contrast to most studies on the subject, the focus here is not on real exchange rates based on aggregate price levels such as the consumer price index (CPI). PPP can only apply to goods and services for which international arbitrage works, that is for tradables. Since still more than half of a typical industrial country's output must be considered nontradable, there is reason to believe that testing PPP on aggregate price indexes will bias the results. The bias introduced by taking nontradables into account may be upward as well as downward, depending on the behaviour of nontradables prices. This may in turn depend on sectoral productivities, per capita income and other factors.<sup>7</sup> Without specific knowledge on these factors, the bias is, of course, difficult to assess and correct afterwards. Consequently, in this study nominal exchange rates are deflated by a measure of the prices of tradables only instead of aggregate price indexes. As about 90 per cent of industrial countries' tradables are made up by manufactured goods, it seems appropriate to approximate the overall indexes of tradables prices by the price indexes for manufactures.

Consistent implicit value-added deflators for manufacturing are obtained from OECD *National Accounts* statistics by dividing the

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<sup>7</sup> This is the Balassa-Samuelson hypothesis. See Balassa (1964) and Samuelson (1964), and for recent empirical applications Summers and Heston (1991) and De Gregorio et al. (1994).

Figure 1 – US Dollar Real Exchange Rates of the Japanese Yen Based on Alternative Deflators, 1972–1995



Source: IMF *International Financial Statistics*, OECD *National Accounts*. Annual averages. CPI and WPI are the consumer and wholesale prices indexes, respectively.

series on manufacturing GDP in current prices by the corresponding series in constant prices.<sup>8</sup> Data on average nominal exchange rates vis-à-vis the US dollar are taken from the IMF *International Financial Statistics* disk. The difference alternative price deflators make for the results regarding long-run PPP can be seen most clearly for the US dollar/yen real exchange rate. Japanese relative nontradables prices have constantly been increasing faster than US relative nontradables prices for the last 30 years due to the faster Japanese productivity growth in tradables (see also Marston 1987). Therefore, one should expect a real appreciation of the yen against the dollar in terms of broad price deflators (with a relative large content of nontradables) but not in terms of tradables prices. Figure 1 presents alternative measures for the real dollar/yen rate. As expected, the measures based on broad price indexes, such as the CPI, the GDP deflator and even the wholesale price index (WPI), exhibit a clear downward trend. In terms of the CPI, the yen has appreciated by 50 per cent since 1973. Of course, such a trend makes it difficult to find real exchange rates to be mean-reverting. The trend is less evident when the real exchange

<sup>8</sup> For Switzerland, sectoral GDP series are not available from the OECD. Tradables prices for Switzerland were instead approximated by the wholesale price index taken from the IMF *International Financial Statistics*.

rate is based on manufacturing deflators; thus, long-run PPP is more likely to hold and the estimated speed of adjustment towards it will be higher.<sup>9</sup>

The analysis is based on annual data. Apparently, the choice of the frequency of the data is not without importance for the results concerning convergence to PPP. Wu (1996) finds that for the same real exchange rates and the same sampling period the *t*-ratio of the autoregression coefficient is the higher (that is, the speed of convergence to PPP is the lower) the lower the frequency of the data is. Sercu et al. (1995) argue that PPP is more likely to hold with low-frequency data. They content that when there are transaction costs for tradables, sluggish adjustment of tradables prices to monetary shocks may cause short-run deviations from PPP that are not eliminated by international arbitrage. This effect disappears over longer sampling intervals when prices have already partly adjusted to monetary shocks so that monetary and price effects tend to coincide. It is therefore less present in low-frequency than in high-frequency data.

As a first step of the formal empirical investigation, single-equation autoregressions were estimated and tests for nonautocorrelation and normality of the residuals were performed. Where necessary to achieve nonautocorrelation, the equation was augmented by lagged first differences of the real exchange rate series. One lagged first difference was sufficient for all but the UK equation for which two lagged first differences were needed. Similar to other studies, nonstationarity of real exchange rates could not be rejected for any of the real exchange rates on the basis of the single-equation results.<sup>10</sup>

The multivariate estimates are found by applying the GLS/SURE estimator to the system of equations determined in the first step. The results are presented in Table 3. The upper part shows for each individual real exchange rate the results of the unrestricted estimation of the system. The *t*-ratios in the middle column of the table are used for inference on nonstationarity. They indicate that the null hypothesis of a random walk is rejected for the real exchange rates of the continental European countries at high levels of significance. Long-run PPP clearly holds for these countries vis-à-vis the United States.

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<sup>9</sup> This pattern is not so clear for some of the continental European countries' US dollar real exchange rates, since in these countries relative nontradables prices only grew faster than in the United States until the beginning of the 1980s.

<sup>10</sup> See for example MacDonald (1995: table 3) and Wu (1996: table 1).

Table 3 – GLS/SURE Test for Nonstationarity of Real Exchange Rates Based on Value-Added Deflators for Manufacturing, 1973–1994<sup>a</sup>

Real exchange rates	$\hat{\rho}$	$t_{\hat{\rho}-1}$	Implied half-life in years (bias-adjusted) <sup>b</sup>
<i>Unrestricted GLS</i>			
Canada	0.73 (0.65)	-2.82 (-2.59)	6.3 (3.6)
France	0.55	-7.34**	2.2
Germany	0.53	-7.77**	2.0
Italy	0.56	-6.71**	1.9
Japan	0.57 (0.70)	-4.03 (-2.00)	2.3 (5.2)
Netherlands	0.57	-7.39**	2.3
Norway	0.68	-7.28**	4.3
Switzerland	0.51	-8.37**	1.8
UK	0.80 (0.76)	-2.82 (-2.15)	15.9 (10.7)
<i>Restricted GLS</i>			
All countries	0.66	-8.65**	3.0
All excl. UK	0.65	-8.54**	2.6
All excl. US and UK <sup>c</sup>	0.51	-9.56**	1.5
Non-EMS <sup>d</sup>	0.67	-6.42**	3.2
All countries (GDP deflator)	0.74	-7.98**	5.2

<sup>a</sup> Model estimated:  $\Delta q_{it} = \mu_i + (\rho_i - 1)q_{it-1} + \sum_{j=1}^p \phi_{ij} \Delta q_{it-j} + \varepsilon_{it}$ ,  $i = 1, \dots, 9$ , on the basis of annual US dollar real exchange rates from 1973 to 1994. \*\* and \* indicate significance at the 5 per cent and the 10 per cent level, respectively, according to the critical values given in Table 1. Figures in parentheses refer to single-equation estimates. – <sup>b</sup> Bias adjustment: for unrestricted GLS:  $\hat{\rho} + 0.173$ , for restricted GLS:  $\hat{\rho} + 0.134$ . – <sup>c</sup> Estimates based on the Deutsche mark real exchange rates of the remaining countries. – <sup>d</sup> Non-EMS: US dollar real exchange rates of Canada, Germany, Japan, Norway, Switzerland and the UK. The 5 per cent critical value simulated for this smaller cross-section is -5.59.

For the UK and Canada,<sup>11</sup> in contrast, nonstationarity cannot be rejected at conventional levels of significance, for Japan only when allowing for a type I error probability of 25 per cent. Convergence towards PPP is consequently not guaranteed. However, the latter result is not to be overinterpreted, given the relative low power of the test with respect to the real exchange rates of these three countries. To increase the power of the tests regarding these parities, the system is re-estimated under the restriction that the autoregression coefficients

<sup>11</sup> The estimates for the Canadian real exchange rates have to be interpreted carefully as the OECD's sectoral output data, on which the price index is based, for this country is not of the same quality as for the other countries.

are equal across equations. As a result, in all equations the  $t$ -ratio increases far above the 5 per cent critical value for rejecting the random-walk hypothesis; nonstationarity is thus also rejected for the UK, Canadian and Japanese parities.

Consider now the speed at which convergence towards PPP takes place. With no cross-equation restriction imposed on the adjustment coefficient  $\rho$ , the point estimates differ considerably. The estimates range from 0.51 for Switzerland to 0.80 (0.76) for the UK.<sup>12</sup> The wide range shrinks, however, when only those real exchange rates are considered which are found stationary. After adding the small-sample bias of 0.173 to each of the  $\hat{\rho}$ 's, the estimates imply that on average a deviation from PPP may take between 1.8 and 4.3 years to have the effects of shock reduced by 50 per cent. The majority of the estimates even lies in the interval of 1.8–2.3 years (22–28 months). Notice that this is considerably shorter than the time periods reported in most previous studies.<sup>13</sup>

For the restricted estimation, the estimated coefficient of 0.66 implies after adjusting for the small-sample bias a half-life of a shock to PPP of 3.0 years, a figure that is still somewhat lower than the consensus estimates cited above. Of course, as this estimate is a weighted average over various adjustment coefficients, its precise value depends on which particular countries' real exchange rates are included in the analysis and on what price measure they are based. Consider the effect of excluding one or two real exchange rates from the analysis. Excluding the real exchange rate with lowest speed towards parity, the British pound, reduces the average half-life estimated by restricted GLS to 2.6 years. Excluding additionally the US dollar and expressing all real exchange rates in relation to the Deutsche mark even reduces the estimated half-life to 1.5 years.<sup>14</sup>

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<sup>12</sup> For the UK, Canada and Japan, single-equation estimates are given in parentheses since due to the low cross-residual correlation between these autoregressions and the rest of the sample, the estimates may be biased. An often encountered rule of thumb is to prefer single-equation before system estimates if the correlation coefficient between the residuals of two equations is smaller than 0.3. According to this criterion, single-equation estimates are to be preferred for the countries above.

<sup>13</sup> Froot et al. (1995) estimate even higher speeds of convergence, albeit for a completely different sample. Examining commodity prices from Holland and England over a period of 700 years they estimate half-lives of deviations from the law of one price of about one year.

<sup>14</sup> In both cases, inference on nonstationarity is qualitatively not affected by excluding two currencies from the sample: Nonstationarity continues to be rejected even for the smaller sample of real exchange rates.

Obviously, convergence towards PPP is even stronger for non-dollar currency pairs which up to now have not received much direct attention in the literature.<sup>15</sup>

It may be argued that the results concerning nonstationarity of real exchange rates are biased by the fact that four of the countries in the sample have been members of a fixed-exchange rate arrangement, the European Monetary System (EMS), during most of the sampling period. To clear this point, the system was re-estimated only for the five non-EMS members plus Germany as a representative of the EMS countries. The result is also given in Table 3. Since the distribution of the test-statistic depends on the number of equations in the system, a new critical value had to be found by simulation. The empirical *t*-ratio clearly exceeds this critical value of  $-5.59$ ; stationarity of the real exchange rates is confirmed. Given the smaller number of relatively fast converging continental European countries in the cross-section, the estimated average speed of convergence is lower than that estimated for the 9-equations system.

Consider, finally, the effect of basing real exchange rates on GDP deflators instead of manufacturing prices when assessing convergence to PPP. The last line of Table 3 shows the result of a restricted estimation of the full system of re-based real exchange rates. The change of the deflator sharply reduces the estimated speed of adjustment: The rate of convergence falls to 14 per cent per year, implying a half-life of a shock of 5.2 years.<sup>16</sup> The use of broad price deflators instead of tradables prices evidently biases upward substantially the estimated rate of convergence to PPP; it, thus, overestimates the swings in a country's external competitiveness and the resulting adjustment efforts.

#### IV. Conclusions

The paper has produced new evidence in favour of the hypothesis that long-run purchasing power parity holds between the industrial

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<sup>15</sup> Note that Wu's (1996) finding of an average half-life of 1.9 years for annual wholesale price-based real exchange rates is perfectly in line with the present results. Wu's sample consists of the real exchange rates vis-à-vis the US dollar of 19 countries, of which, apart from Canada and Japan, all are European. Presumably, most of these currencies will converge to PPP with the dollar at a rate similar to that estimated for the continental European countries, implying half-lives of a shock of about 2 years. Given the large number of these fast-convergers in Wu's study, it is not surprising that his estimated average is also close to 2 years.

<sup>16</sup> This figure corresponds closely to that found by Frankel and Rose (1996) for a similar sample. See footnote 3 above.

countries. A panel test was employed which uses more efficiently the available time-series information on floating exchange rates than the usual single-equation procedure but which still allows to assess the rate of convergence to PPP for each pair of currencies separately. In accord with other studies, the paper finds that convergence to PPP is in fact present. Rates of convergence to PPP, however, seem to be quite different across countries. For most of the OECD countries, the average rate of convergence to PPP appears to be higher than previously found. The average point estimate for the degree of mean-reversion for these countries' US dollar real exchange rates is 0.70, implying that a shock to PPP is reduced by 50 per cent after 2 years. There is also evidence that long-run PPP between these countries is attained even faster. For other currencies, however, notably the British pound and the Japanese yen, PPP deviations vis-à-vis the dollar have proved to be more persistent. The influence of these slowly converging parities together with the use of broad price indexes in the construction of real exchange rates largely explains why previous panel studies found convergence to be slower.

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**Abstract:** Assessing Convergence to Purchasing Power Parity: A Panel Study for 10 OECD Countries. – The paper analyzes whether post-Bretton Woods real exchange rates of ten OECD countries are nonstationary so that long-run purchasing power parity (PPP) can be considered to hold. A test procedure is proposed which treats the various real exchange rates as a panel but still allows to assess the rate of convergence to PPP for each pair of currencies separately. It is shown that long-run (tradables-) PPP holds between all countries analyzed. Rates of convergence to PPP, however, are found to be quite different across countries. For most of the OECD countries convergence takes place faster than previously indicated. JEL no. F31



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**Zusammenfassung:** Gibt es eine Konvergenz zur Kaufparität? Eine Panel-Untersuchung für 10 OECD-Länder. – Der Verfasser untersucht, ob die realen Wechselkurse nach Bretton Woods für 10 OECD-Länder nichtstationär sind, so daß die Hypothese einer langfristigen Kaufkraftparität (PPP) als haltbar angesehen werden kann. Für den Test wird ein Vorgehen empfohlen, bei dem die verschiedenen realen Wechselkurse als ein Panel behandelt werden, das es aber trotzdem erlaubt, die Konvergenzrate zur PPP für jedes Länderpaar getrennt zu beurteilen. Es zeigt sich, daß die Hypothese einer langfristigen PPP (für handelbare Güter) zwischen allen analysierten Ländern haltbar ist. Allerdings erweisen sich die Konvergenzraten zur PPP als ganz verschieden von Land zu Land. Für die meisten OECD-Länder erfolgt die Konvergenz schneller als bisher angenommen.

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