Real Exchange Rates and Unit Root Tests

By

Ashok Parikh and **Elizabeth Wakerly**

Contents: I. Introduction. – II. Bilateral and Multilateral Real Exchange Rates. – III. Unit Root Tests of Real Exchange Rates. – IV. Panel Data Unit Root Tests. – V. Cointegration between Nominal Exchange Rates and Relative Prices. – VI. Summary and Conclusions. – Appendix.

I. Introduction

This paper argues that real exchange rates should not have unit roots: where they do, it implies violation of the strict version of purchasing power parity (PPP). However, it does not imply rejection of a modified version of PPP based on cointegration between nominal exchange rates and relative prices.

The standard approach to unit root testing, where the null of stationarity is tested against the alternative of non-stationarity, leads to results that tend to favour the null. Changing the null to non-stationarity, however, again leads to results that tend to favour the null, this time supporting non-stationarity. These contradictory findings are largely due to the low power of unit root tests when roots are close to unity in absolute value.[1]

Roll (1979) argues that a random walk is a sensible null hypothesis, since real exchange rate changes are similar to changes in asset prices and as such should not be predictable. An alternative rationale for this null comes from the literature on productivity innovations: if these are permanent then real exchange rates may be non-stationary. Blough (1992) objects to conducting unit root tests at all, since the level and maximum power of generic unit root tests must be equal. Without prior

Remark: The authors would like to thank an anonymous referee for helpful comments in improving the paper.

[1] Christiano and Eichenbaum (1990), Cochrane (1991) and Blough (1992) show that any trend-stationary process can be approximated in finite samples by a unit root process (and vice versa), implying that any test of one against the other must have power no greater than size. If the autoregressive parameter is close to unity, its estimate will have a normal asymptotic distribution. On the other hand, the unit root asymptotic distribution provides a better finite sample approximation, and hence the investigators could err in favour of a unit root. In an analogous manner, near-unit root variables are better forecast by unit root models than stationary models.

information, the test cannot be informative in distinguishing between unit root and stationary processes (classes of unit root tests and stationary processes could be observationally equivalent).

Huizinga (1987) finds that real exchange rates exhibit a random walk: individual series do not have stable means and any innovation has permanent effects. Most pre-float studies, however, support the existence of stable exchange rates over the long run. In particular, the monetary approach to exchange rate determination adopted the strong proposition of PPP. In the 1980s, most studies rejected PPP: nominal exchange rates were not cointegrated with relative prices (where the coefficient of relative prices was assumed to be unity).

Recent univariate studies of EMS exchange rates (Zivot and Andrews 1994; Taylor 1995; Rogoff 1996) report ambiguous results, but tend to favour rejecting non-stationarity. When using long-run data (for two centuries), Lothian and Taylor (1996) show that strict PPP normally holds. Studies conducting unit root tests on panel exchange rate data (Coakley and Fuertes 1997; MacDonald 1996; Wu 1996; Oh 1996) tend to reject the null of a unit root in favour of mean reversion (although different tests produce different results).

We suggest that where economic phenomena (such as exchange rates) are based on aggregate behaviour, the pattern of such behaviour can often be predicted by some economic theory or law, allowing restrictions to be imposed *a priori*. This might permit the researcher to assume, for example, stationarity *a priori*, avoiding the need for unit root tests altogether.

The economic argument against unit roots in real exchange rates is strong. Multilateral trade-weighted real exchange rates ought to exhibit stationarity. Although individual countries could be in disequilibrium for long periods, it is unlikely that multilateral real exchange rates could sustain disequilibrium for very long: relative prices cannot sustain a stochastic trend and will eventually revert to the mean of the series. In the UK, for example, oil prices (relative to the general price level) exhibited a stochastic trend in 1973–74 and 1978–79, but by the late 1980s the relative price of oil was little changed from its pre-1974 level.[2]

This paper examines monthly OECD exchange rate data (1979–1997) using various unit root tests.[3] We show that most of these tests

[2] In real terms, energy prices were lower in the late 1980s than they were before the oil price shocks.

[3] Levin and Lin (1993) and Im et al. (1998) have recently proposed panel unit root tests, designed to increase sample size and improve the asymptotic power of the test.

support the hypothesis of a unit root. But tests of cointegration between nominal exchange rates and relative prices, based on a likelihood ratio (LR) test (see Larsson et al. 1998), reject the null hypothesis of a unit root.

We suggest that researchers need not conduct unit root tests on real exchange rate data (i) when a modified version of PPP is used; or (ii) if there is a long enough time series (50 years). Given the definition of real exchange rates, the indicator should be stationary and should have intrinsic mean reverting behaviour.

In Section II of this paper we define bilateral and multilateral real exchange rates. In Section III we conduct unit root tests on bilateral and multilateral real exchange rates of individual countries. In Section IV we outline the procedures for the panel unit root tests. In Section V we conduct the panel cointegration tests, and in Section VI we provide a summary and conclusions.

II. Bilateral and Multilateral Real Exchange Rates

Monthly exchange rate data for 16 countries are available from the IMF/IFS CD-ROM database for the period January 1979–December 1997. The Exchange Rate Mechanism (ERM) of the EMS has influenced the evolution of exchange rates for many countries in our sample period since 1979; for this reason we exclude data prior to 1979.[4]

All exchange rates are expressed in dollars per unit of currency. Real exchange rates are derived using both producer and consumer price indices for each country (REX1 and REX2, respectively):

$$REX1_t = -\log E_t - \log P^*_{pt} + \log P_{pt}, \tag{1}$$

$$REX2_t = -\log E_t - \log P^*_{ct} + \log P_{ct}, \tag{2}$$

[4] The formal objective of the EMS has been the stabilization of member countries' nominal exchange rates within generally narrow pre-agreed bounds. This limit was set at ± 2.25 per cent from central parity for Belgium, France, Germany and Netherlands in 1979. For the Italian lira the limit was ± 6 per cent, which was changed to ± 2.25% from January 1990. As well as the provision of a margin of fluctuation, realignments were permitted. In September 1992, speculative pressures against some member countries forced Italy and the UK to withdraw from the ERM. The Portuguese escudo and Spanish peseta were devalued several times throughout 1992 and 1993, and in February 1993, the Irish punt was also devalued. The exchange rate crises ended in August 1993 with the expansion of bands from ± 2.25 to ± 15 per cent for all bilateral rates with the exception of the guilder/DM rate. The perception is that wider bands have reduced the prospects for exchange rate crises in the ERM. In March 1995, however, those countries within the ERM again experienced sharp fluctuations and the escudo and peseta were again devalued.

where E_t is the nominal exchange rate (dollars per unit of local currency), P^*_{pt} (P^*_{ct}) is the time t producer (consumer) price index for a foreign country, and P_{pt} (P_{ct}) is the US producer (consumer) price index. Consumer prices are most frequently used as indicators of inflation since they reflect changes in the costs of a fixed basket of goods for the average consumer. However, it is likely *a priori* that PPP will be violated more often for real exchange rate data based on consumer prices than for real exchange rate data based on producer prices, since the former include taxes and retail margins which differ across countries.

Multilateral indices generally give a more accurate picture of a country's international competitiveness and are therefore to be preferred over bilateral indices that are most commonly calculated against the dollar. The IMF has compiled real effective exchange rates for industrial countries, and for 17 of these, country data on normalized unit labour costs in manufacturing are also available.[5] A multilateral exchange rate index (1990=100) is compiled using weighted aggregate data for trade in manufactured goods, averaged over the period 1989–1991. Weights reflect both the relative importance of a country's trading partners in direct bilateral relations and that resulting from competition in third markets. In all bilateral and multilateral definitions used here, an increase in an index represents a real exchange rate appreciation or a deterioration in a country's degree of competitiveness; conversely for a decrease in the index.

III. Unit Root Tests of Real Exchange Rates

We investigate the long-run mean reversion properties of the real exchange rate series for each OECD country. Table A1 in the Appendix contains results of three unit root tests (each with the null hypothesis of a unit root). The equation used to test for a random walk has the following form:

$$REXi_t = \alpha_0 + \alpha_1 t + seasonals + \alpha_2 REXi_{t-1} + \phi(l)\, REXi_t + \varepsilon_t, \quad (3)$$

where i refers to country i's real exchange rate vis-à-vis the dollar, l is the lag operator, $\phi(l)$ is a pth order polynomial in l with coefficients ϕ_1, ϕ_2, ϕ_3, ... ϕ_p, and ε_t is a white noise error term. Under the null hy-

[5] The nominal effective exchange rate index represents the ratio of an index of the period average exchange rate of the currency in question to a weighted average of exchange rates for the currencies of selected countries. A real effective exchange rate index is adjusted for relative movements in national prices or cost indicators of the home country and selected countries.

pothesis, REX has a unit root, i.e. $\alpha_2 = 1$; under the alternative, PPP holds in the long run and the real exchange rate is mean reverting so that $\alpha_1 = 0$ and $\alpha_2 < 1$.

We use a standard augmented Dickey-Fuller (ADF) test, the Phillips-Perron (PP) test, and the weighted symmetric (WS) test (Pantula et al. 1994). In each case, the optimal lag length is chosen using the Akaike Information Criterion (AIC). The results of these tests support the null hypothesis of non-stationarity.

We also test the hypothesis that the principle of PPP does not admit the presence of a deterministic trend or multiple unit roots (Enders 1995), by restricting α_1 in (3) to equal zero. The estimated equation now becomes:

$$REXi_t = \alpha_0^* + \alpha_2^* REXi_{t-1} + \phi(l)\, REXi_t + \varepsilon_t \tag{4}$$

with the null $\alpha_2^* = 1$ tested against the alternative $\alpha_2^* < 1$. The results are presented as the tests without trend. We cannot reject the null of non-stationarity.

Multilateral exchange rates are also examined country by country using the above tests. Once again, the results all tend to favour the null of non-stationarity.

IV. Panel Data Unit Root Tests

Im et al. (IPS) (1998) propose two new tests for panel unit root study based on the following equation:

$$\Delta REXi_t = \alpha_i + \beta_i REXi_{t-1} + \phi(l)\, \Delta REXi_t + \varepsilon i_t . \tag{5}$$

The standardized t-bar statistic is derived from the average t-ratio for testing the null of a unit root, $\beta_i = 0$ for all i.[6] The test statistic is given by:

$$\Lambda_i = \frac{\sqrt{N}(T)\,[\bar{t}_T - E(t_T)]}{\sqrt{Var(t_T)}}, \tag{6}$$

where $E(t_T)$ and $Var(t_T)$ are the asymptotic values of the mean and variance, respectively, of the average ADF statistic (tabulated by IPS for various lag lengths and sample sizes). Where the ADF regression has different augmentation lags (L_i) for each country, the terms $E(t_T)$ and

[6] The average β_i used in the IPS test differs from the long-run cointegration coefficient as defined by Phillips and Moon (1999).

$Var(t_T)$ in (6) are replaced by the group averages of the tabulated values of $E(t_T, L_i)$ and $Var(t_T, L_i)$, respectively. The standardized LM-bar statistic uses the average LM value:[7]

$$\Gamma_{\overline{LM}} = \frac{\sqrt{N}[\overline{LM}_{NT} - E(\eta_T)]}{\sqrt{Var(\eta_T)}}. \tag{7}$$

Under the null of a unit root, both the t-bar and LM-bar statistics follow the standard normal distribution for $N \to \infty$, $T \to \infty$ and $N/T \to 0$. Simulations conducted by IPS show that both tests have substantially more power than standard ADF tests and the Levin-Lin (Levin and Lin 1992) panel unit root test.[8]

Both tests are also applied to demeaned data where a time-specific common effect allows for a degree of dependency across groups: cross-section means are subtracted from the original data and the statistics presented in (6) and (7) are calculated.

A strict interpretation of PPP requires a constant long-run real exchange rate. This interpretation was used in testing for unit roots in the individual country equations (1) and (2). Here, we continue to impose symmetry, but allow the cointegrating vector coefficients to deviate from unity:[9]

$$\log E_t = \mu_0 + \mu_1(\log P_{pt} - \log P^*_{pt}) \tag{8}$$

$$\log E_t = \nu_0 + \nu_1(\log P_{ct} - \log P^*_{ct}). \tag{9}$$

Equations (8) and (9) are used to predict nominal exchange rates. Real exchange rates are then derived for producer and consumer prices, respectively, with the coefficient on relative prices estimated μ_i or ν_i $(i = 1, \ldots, 15)$ (producer and consumer prices, respectively) for each country. This PPP-predicted real exchange rate series is also tested for the presence of a unit root.[10]

Larsson et al. (1998) derive a likelihood-based cointegration test for heterogeneous panels which is more powerful than the rank test for in-

[7] Expected values and variances are tabulated by IPS.
[8] Although the IPS test itself may be an inefficient estimator as it fails to correct for cross-sectional dependence in real exchange rates (O'Connell 1998).
[9] This approach can be justified given the problems of measurement error and non-identical consumer baskets (Cheung and Lai 1993).
[10] Results for panel unit root tests on both real and predicted real exchange rates are available from the authors on request. In most cases, the t-bar tests suggest that a unit root cannot be rejected for the real exchange rate data, but can be rejected for the predicted real exchange rate series. None of the series reject a unit root when the LM-bar test is used.

dividual countries. The LR-bar test statistic is based on the average trace value derived from individual country trace statistics for different cointegrating ranks:

$$\begin{aligned} &\textit{Likelihood average trace value}\,[H(r)/(H(p)] \\ &\qquad = \frac{1}{N}[\sum LR_{it}(H(r)/H(p))] \end{aligned} \tag{10}$$

$$\gamma_{\overline{LR}[H(r)/H(p)]} = \frac{\sqrt{N}\,[LR_{NT}(H(r)/H(p)) - E(Z_k)]}{\sqrt{Var}\,(Z_k)}, \tag{11}$$

where $E(Z_k)$ and $Var(Z_k)$ are the moments estimated for different values of $(p - r)$ where p is the number of variables and r is the cointegrating rank.

V. Cointegration between Nominal Exchange Rates and Relative Prices

We estimate the model given by (8) and (9) using the Johansen and Juselius (1990) maximum likelihood approach and apply trace tests at a country level. Lags are selected using the Akaike Information Criterion (AIC). Tables 1 and 2 present the results.

The null hypothesis of zero cointegrating vectors is rejected at the 1 and 5 per cent levels, respectively, for producer price-based and consumer price-based exchange rates. Individual trace tests on the consumer price-based data suggest that the null can be rejected in favour of one cointegrating relationship for seven countries.

We reject the null hypothesis of no cointegrated relationship in favour of one.[11] The results suggest that there is one cointegrating relationship in the two-variable system: a cointegrating relationship between nominal exchange rates and the relative price index for each country against the USA.

VI. Summary and Conclusions

This paper argues that unit root tests have low power against the alternative of near-unit root stationary time series. We suggest that *a priori* information, such as that provided by an economic law or theory, should

[11] Given the short time series, the power of the panel rank test is low, increasing the probability of choosing a rank higher than one. We regard the panel rank test result which rejects one cointegrating vector in favour of two as suspect.

Table 1 – *Trace Test for Cointegration: Nominal Exchange Rates and Relative Producer Price Indices*[a]

Country	Lag (k_i)	LR_{it} [$H(r) \mid H(2)$]		Rank (5%)	Rank (1%)
		$r = 0$	$r = 1$	r_i	r_i
Australia	2	20.23	2.10	1	0
Austria	8	26.40	0.66	1	1
Belgium	4	23.82	2.66	1	1
Canada	2	18.08	2.45	1	0
Denmark	1	28.03	6.48	2	1
Finland	1	16.02	4.04	0	0
Germany	1	28.53	2.98	1	1
Greece	1	19.61	5.31	1	0
Ireland	1	18.36	6.38	1	0
Italy	4	21.94	4.36	1	1
Japan	1	6.52	0.33	0	0
Netherlands	1	31.28	3.05	1	1
Spain	1	36.00	4.72	1	1
Sweden	1	21.22	3.36	1	1
United Kingdom	1	10.69	2.72	0	0
$\gamma_{\overline{LR}}$	5%	18.73 [b]	5.998 [c]	n/a	n/a

[a] For country-by-country tests, the critical values of the rank test at the 5 per cent level are 17.84 and 8.083, respectively, for testing $r = 0$ and $r = 1$, while at 1 per cent the corresponding values are 21.96 and 11.57. The panel rank test has the critical value 1.645. – [b] Reject the null hypothesis of zero cointegrating vectors. – [c] Reject the null hypothesis of one cointegrating vector.

be used where available. If this implies restrictions much stronger than a unit root, then the hypothesis of stationarity should be a maintained hypothesis, perhaps requiring no other tests.

Various investigators have queried the power of unit root tests when applied to a short sample. New studies favour panel data unit root tests such as the t-bar, LM-bar and LR-bar tests. Using these tests, we reject the unit root hypothesis for OECD multilateral real exchange rate data and predicted real exchange rates. We believe, however, that real exchange rates cannot sustain a long-run stochastic trend; if we had found a unit root it would simply have been an artefact of the data. The economic law of PPP seems to hold true on long time series data, and hence, for a suitably long time series, real exchange rates should be stationary.

Table 2 – *Trace Test for Cointegration: Nominal Exchange Rate and Relative Consumer Price Indices*[a]

Country	Lag (k_i)	LR_{it} [$H(r) \mid H(2)$]		Rank (5%)
		$r = 0$	$r = 1$	r_i
Australia	12	13.52	0.54	0
Belgium	12	16.72	1.95	1
Canada	1	2.43	0.24	0
Denmark	12	20.03	7.81	1
Finland	2	8.67	3.17	0
France	3	27.84	6.13	1
Germany	1	31.13	2.00	1
Greece	12	14.67	4.31	0
Italy	1	31.68	3.99	1
Japan	12	20.82	2.35	1
Netherlands	12	12.54	1.64	0
Portugal	1	30.81	4.80	1
Spain	12	14.27	5.00	0
Sweden	1	8.60	2.72	0
United Kingdom	12	10.88	0.84	0
$\gamma_{\overline{LR}}$	5%	13.79[b]	5.284[c]	

[a] For country-by-country tests, the 5 per cent critical values of the rank test are 17.84 and 8.083, respectively, for testing $r = 0$ and $r = 1$. The panel rank test has the critical value 1.645. Excluding Canada, Finland and Sweden from calculation of the panel rank test statistic gives values for $\gamma_{\overline{LR}}$ of 15.29 and 5.380 for $r = 0$ and $r = 1$, respectively. – [b] Reject the null hypothesis of zero cointegrating vectors. – [c] Reject the null hypothesis of one cointegrating vector.

Appendix

Table A1 – *Unit Root Tests of Bilateral and Multilateral Real Exchange Rates*[a]

Country		Weighted symmetric		Augmented Dickey-Fuller		Phillips-Perron		Definition
		with trend	without trend	with trend	without trend	with trend	without trend	
Australia	RAUI (3)	−2.11	−2.31	−2.14	−2.08	−9.30	−8.34	Producer price
	RAUI1	Data not available						
	LAUI (3)	−2.61	−1.56	−2.50	−1.41	−11.58	−5.30	Multilateral
Austria	ROEI (3)	−1.58	−1.87	−2.06	−1.74	−7.55	−5.57	Producer price
	ROEI1 (3)	−1.42	−1.77	−2.16	−1.64	−6.47	−4.36	Consumer price
	LOEI (5)	−1.30	0.04	−2.62	0.42	−10.52	0.74	Multilateral
Belgium	RBGI (3)	−1.34	−1.52	−2.40	−2.25	−8.53	−7.94	Producer price
	RBGI1 (5)	−1.05	−1.33	−2.24	−1.94	−6.63	−5.85	Consumer price
	LBGI (3)	−0.36	0.24	−2.53	−3.14	−4.51	−5.82	Multilateral
Canada	RCNI (12)	−2.17	−2.12	−1.91	−1.94	−5.08	−5.51	Producer price
	RCNI1 (12)	−1.99	−1.76	−1.69	−1.50	−2.48	−1.84	Consumer price
	LCNI (8)	−1.37	−1.36	−1.43	−1.43	−4.55	−4.64	Multilateral
Denmark	RDKI (3)	−1.42	−1.75	−2.15	−1.56	−7.12	−4.19	Producer price
	RDKI1 (3)	−1.24	−1.62	−2.19	−1.74	−6.59	−4.72	Consumer price
	LDKI (3)	−1.44	−1.25	−3.45[b]	−0.84	−11.46	−1.28	Multilateral
France	RFRI	Data not available						
	RFRI1 (5)	−1.54	−1.75	−2.13	−1.91	−6.62	−5.82	Consumer price
	LFRI (11)	−2.88	−0.88	−2.85	−0.64	−18.73	−3.32	Multilateral
Finland	RFNI (3)	−1.97	−2.03	−1.70	−1.76	−5.54	−5.57	Producer price
	RFNI1 (3)	−1.78	−1.88	−1.61	−1.65	−4.67	−4.72	Consumer price
	LFNI (3)	−0.95	−0.54	−1.86	−0.00	−5.05	0.21	Multilateral
Germany	RBDI (3)	−1.35	−1.73	−2.07	−1.51	−6.98	−4.02	Producer price
	RBDI1 (3)	−1.25	−1.61	−2.29	−1.86	−7.00	−5.29	Consumer price
	LBDI (4)	−1.19	−0.97	−2.38	−0.64	−10.09	−0.68	Multilateral
Greece	RGRI (3)	−1.54	−1.72	−2.13	−1.55	−7.03	−4.59	Producer price
	RGRI1 (12)	−1.86	−1.96	−2.99	−2.01	−6.44	−4.21	Consumer price
	LGRI	Data not available						

Table A1 – Continued

Country		Weighted symmetric		Augmented Dickey-Fuller		Phillips-Perron		Definition
		with trend	without trend	with trend	without trend	with trend	without trend	
Ireland	RIRI (3)	–1.82	–1.98	–2.28	–1.92	–7.09	–5.62	Producer price
	RIRI1	Data not available						
	LIRI (3)	–1.23	1.15	–2.70	1.20	–10.65	1.11	Multilateral
Italy	RITI (3)	–1.85	–1.90	–1.96	–1.58	–5.95	–3.95	Producer price
	RITI1 (3)	–1.93	–2.03	–1.87	–1.75	–5.55	–4.86	Consumer price
	LITI (7)	–1.83	–1.68	–1.72	–1.36	–6.14	–3.95	Multilateral
Japan	RJPI (3)	–1.47	–1.81	–1.70	–1.49	–7.33	–3.69	Producer price
	RJPI1 (5)	–1.22	–1.72	–1.95	–1.36	–9.08	–3.36	Consumer price
	LJPI (4)	–1.22	–1.72	–2.25	–1.36	–11.89	–3.36	Multilateral
Netherlands	RNLI (3)	–1.23	–1.67	–2.17	–1.51	–7.46	–4.08	Producer price
	RNLI1 (3)	–1.35	–1.58	–2.35	–2.11	–7.40	–6.56	Consumer price
	LNLI (12)	–1.35	1.51	–4.74	–1.90	–12.91	–5.76	Multilateral
Portugal	RPII	Data not available						
	RPII1 (5)	–1.45	–1.71	–2.01	–1.30	–6.14	–3.04	Consumer price
	LPII (5)	–1.96	–0.87	–1.82	–0.81	–6.55	–2.03	Multilateral
Spain	RESI (5)	–1.67	–1.85	–1.98	–1.76	–4.78	–4.27	Producer price
	RESI1 (5)	–1.48	–1.72	–1.98	–1.65	–4.34	–3.54	Consumer price
	LES1 (5)	–1.65	–1.71	–2.17	–1.99	–5.38	–5.41	Multilateral
Sweden	RSDI (3)	–1.73	–1.89	–1.90	–1.67	–6.06	–5.13	Producer price
	RSDI1 (3)	–1.48	–1.57	–1.76	–1.72	–5.10	–4.97	Consumer price
	LSDI (3)	–2.24	–1.29	–2.00	–1.45	–6.37	–3.50	Multilateral
United Kingdom	RUKI (5)	–2.24	–1.70	–2.18	–1.42	–8.43	–9.95	Producer price
	RUKI1 (5)	–2.21	–2.18	–2.04	–1.97	–7.36	–7.16	Consumer price
	LUKI (5)	–1.14	–1.82	–2.26	–1.97	–14.45	–10.26	Multilateral
United States	LUSI (3)	–0.63	–1.52	–1.78	–1.11	–5.37	–2.16	Multilateral

[a] Similar tests were conducted on predicted bilateral real exchange rates based on producer and consumer prices and also on residuals of a cointegrating regression between nominal exchange rates and relative prices. The results based on residuals suggested rejection of the null hypothesis of a unit root. All other results tended to confirm non-stationarity. – [b] Reject H_0 at the 5 per cent level.

References

Blough, S. R. (1992). The Relationship between Power and Level for Generic Unit Root Tests in Finite Samples. *Journal of Applied Econometrics* 7 (3): 295–308.

Christiano, L. J., and M. Eichenbaum (1990). Unit Roots in Real GNP: Do We Know and Do We Care? *Carnegie-Rochester Conference Series on Public Policy* 32: 7–61.

Coakley, J., and A. M. Fuertes (1997). New Panel Unit Root Tests of PPP. *Economics Letters* 57 (1): 17–22.

Cochrane, J. (1991). A Critique of the Application of Unit Root Tests. *Journal of Economic Dynamics and Control* 15 (2): 275–284.

Cheung, Y., and K. S. Lai (1993). Finite Sample Sizes of Johansen's Likelihood Ratio Tests for Cointegration. *Oxford Bulletin of Economics and Statistics* 55 (3): 313–328.

Enders, W. (1995). *Applied Econometric Time Series.* New York: Wiley.

Huizinga, J. (1987). An Empirical Investigation of the Long-Run Behaviour of Real Exchange Rates. *Carnegie-Rochester Conference Series on Public Policy* 27: 149–214.

Im, K. S., M. H. Pesaran, and Y. Shin (1998). Testing for Unit Roots in Heterogeneous Panels. University of Cambridge mimeo, Cambridge.

Johansen, S., and K. Juselius (1990). Maximum Likelihood Estimation and Inference on Cointegration – with Applications to the Demand for Money. *Oxford Bulletin of Economics and Statistics* 52 (2): 169–210.

Larsson, R., J. Lyhagen, and M. Lothgren (1998). Likelihood Based Cointegration Tests in Heterogeneous Panels. Working Paper Series in Economics and Finance No. 250. Stockholm School of Economics, Stockholm.

Levin, A., and C-F. Lin (1992). Unit Root Tests in Panel Data: Asymptotic and Finite-Sample Properties. University of California Working Paper 92, 23. University of California, San Diego.

Lothian, J., and M. Taylor (1996). Real Exchange Rate Behaviour: The Recent Float from the Perspective of the Past Two Centuries. *Journal of Political Economy* 104 (3): 488–509.

MacDonald, R. (1996). Panel Unit Root Tests and Real Exchange Rates. *Economics Letters* 50 (1): 7–11.

O'Connell, P. G. J. (1998). The Overvaluation of Purchasing Power Parity. *Journal of International Economics* 44 (1): 1–19.

Oh, K. Y. (1996). Purchasing Power Parity and Unit Root Tests using Panel Data. *Journal of International Money and Finance* 15 (3): 405–418.

Pantula, S. G., G. Gonzalez-Farias, and W. A. Fuller (1994). A Comparison of Unit-Root Test Criteria. *Journal of Business and Economic Statistics* 12 (4): 449–459.

Phillips, P. C. B., and H. R. Moon (1999). Nonstationary Panel Data Analysis: An Overview of Some Recent Developments. Cowles Discussion Paper 1221. Yale University, New Haven.

Rogoff, K. (1996). The Purchasing Power Parity Puzzle. *Journal of Economic Literature* 34 (2): 647–668.

Roll, R. (1979). Violations of Purchasing Power Parity Theory and their Implications for Efficient International Commodity Markets. In M. Sarnat and G. Szego (eds.), *International Finance and Trade.* Cambridge, Mass.: Ballinger.

Taylor, M. (1995). The Economics of Exchange Rates. *Journal of Economic Literature* 33 (1): 13–47.

Wu, Y. (1996). Are Real Exchange Rates Non-Stationary? Evidence from a Panel Data Test. *Journal of Money, Credit and Banking* 28 (1): 54–63.

Zivot, E., and D. W. K. Andrews (1992). Further Evidence on the Great Crash, the Oil Price Shock and the Unit Root Hypothesis. *Journal of Business and Economic Statistics* 10 (3): 251–270.

***

Abstract: Real Exchange Rates and Unit Root Tests. – This paper examines monthly OECD exchange rate data (1979–1997) using univariate and panel data unit root tests. Some of these tests support the hypothesis of a unit root. But tests of cointegration reveal the existence of weak purchasing power parity relationships between bilateral nominal exchange rates and relative prices. We suggest that researchers need not conduct unit root tests on real exchange rate data when a modified version of PPP is used; or if there is a long enough time series. Given the definition of real exchange rates, the indicator should be stationary and should have intrinsic mean reverting behaviour. JEL no. C23, F31

*

Zusammenfassung: Reale Wechselkurse und Tests auf Einheitswurzel. – Dieser Artikel untersucht monatliche OECD-Daten über Wechselkurse in den Jahren 1979–1997, wobei die Tests eindimensional sind und Paneldaten verwenden. Einige der Tests stützen die Hypothese der Einheitswurzel. Aber Kointegrationstests zeigen, dass nur schwache Kaufkraftparitätenbeziehungen zwischen bilateralen nominalen Wechselkursen und relativen Preisen bestehen. Anderen Forschern wird vorgeschlagen, die realen Wechselkurse keinem Test zu unterwerfen, wenn eine modifizierte Version der Kaufkraftparitätentheorie verwendet wird oder wenn die Zeitreihe lang genug ist. Bei gegebener Definition der realen Wechselkurse sollte der Indikator stationär sein und von sich aus zum Mittelwert zurückkehren.