A Simple Test for PPP Among Traded Goods*

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Abstract

The so-called Balassa-Samuelson model implies that relative prices of non-traded goods may be nonstationary and, hence, that PPP should preferably be tested on real exchange rates based on prices of traded goods only. We propose a simple test for PPP among traded goods which can be applied to real exchange rates based on prices of all (that is, both traded and non-traded) goods. We show through simulations that the test is reliable for a sample size commonly considered in practice. Upon applying the test to bilateral real exchange rates based on the general CPI among a group of industrialized countries during the recent float, we find little evidence in favor of PPP among traded goods. This does not change when we use real exchange rates based on various components of the CPI.

Keywords: purchasing power parity; unit roots; unidentified nuisance parameters.
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1 Introduction

Purchasing Power Parity [PPP] is generally believed to hold as a long-run equilibrium relationship. However, the empirical support for PPP, when examined by unit root tests, is not overwhelming[1]. In particular, such unit root tests typically suggest nonstationarity of real exchange rates when applied to data from the post-Bretton Woods era. This finding often is attributed to the lack of power of standard unit root and cointegration tests in small samples and/or against certain near-nonstationary alternatives. To surmount this lack of power, several different strategies have been devised, such as the use of long spans of data[2] and the use of panel unit root tests[3]. Although these approaches lead to more evidence supporting PPP, they are not without criticism. The use of long spans of data inevitably means mixing data from fixed and floating (nominal) exchange rate regimes. It seems likely that real exchange rates also behave differently under such different exchange rate regimes, see Grilli and Kaminsky (1991) for an elaborate discussion. O'Connell (1998) criticizes the results from studies which employ panel unit root tests for not correcting for the comovements of the real exchange rates in the panels used. Taking this into account greatly reduces the statistical support for PPP, see also Pedroni (2001) and Wu and Wu (2001). Additionally, Sarno and Taylor (1998) and Taylor and Sarno (1998) argue that panel unit root tests may be misleading as rejection of the null hypothesis (which is non-stationarity of *all* real exchange rates in the panel) may occur if only a small number of the series considered actually is stationary.

Next to statistical arguments, there are also economic explanations for the failure to find empirical evidence in favor of PPP. Currently, the most popular among these is the existence of barriers to trade, such as tariffs and transportation costs. Several theoretical models have been developed showing that such market frictions lead to nonlinear adjustment in real exchange rates to the PPP equilibrium rate, see Benninga and Protopapadakis (1988), Dumas (1992) and Sercu, Uppal and Van Hulle (1995), among others. In particular, the strength of mean reversion increases with the deviation from parity, as the profits from goods arbitrage, which is generally thought to be the ultimate force behind maintaining PPP, do not make up for the costs involved in the necessary transactions for small deviations from the presumed

[1] Comprehensive surveys of the voluminous literature on PPP can be found in Froot and Rogoff (1995), Rogoff (1996), and, more recently, Sarno and Taylor (2001).

[2] See Abuaf and Jorion (1990), Ardeni and Lubian (1991), Edison (1987), Glen (1992), Grilli and Kaminsky (1991), and Lothian and Taylor (1996), among others.

[3] See Frankel and Rose (1996), Lothian (1997), MacDonald (1996), Oh (1996), Papell (1997), Papell and Theodoridis (1998), Taylor and Sarno (1998) and Wu (1996), among others.

equilibrium real exchange rate. This has recently led to an outburst of empirical papers using nonlinear time series models to accommodate such effects[4].

Another well-known and intuitively plausible economic explanation for the lack of empirical support for PPP is the presence of non-traded goods in the price indices used to construct real exchange rates. Given that PPP is supposed to be maintained by international commodity arbitrage, it follows immediately that this applies to traded goods only. Moreover, as already shown by Balassa (1964) and Samuelson (1964), permanent productivity shocks can lead to permanent changes in the relative prices of traded and non-traded goods, which in turn implies that the real exchange rate has a nonstationary component. The empirical evidence for the Balassa-Samuelson effect is mixed, see Canzoneri, Cumby and Diba (1999), De Gregorio, Giovannini and Wolf (1994), and Heston, Nuxoll and Summers (1994) for some favorable results. Engel (1999) and Rogers and Jenkins (1995) document, however, that changes in relative prices of non-traded goods appear to account for only a small fraction of real exchange rate changes, suggesting that the Balassa-Samuelson is not a very important factor in explaining the empirical failure of PPP. Moreover, Engel (2000) shows that, under these circumstances, standard unit root tests suffer from large size distortions, especially when applied to long historical time series, see also Ng and Perron (2001). Note that this finding actually makes it even more puzzling that the support for PPP is rather scarce.

In the Balassa-Samuelson model, PPP is assumed to hold among traded goods due to the forces of international commodity arbitrage, such that the relative prices of traded goods are stationary and mean-reverting. Several studies show, however, that PPP may fail even among traded goods, see Canzoneri *et al.* (1999), Engel (1993), Engel and Rogers (1996, 2001), and Parsley and Wei (2001), among others. Apart from the impediments to trade mentioned before, another explanation for this finding is local-currency-pricing [LCP] or pricing-to-market [PTM], meaning to say that producers selling abroad set prices in the currency of consumers rather than their own. Under LCP, changes in nominal exchange rates do not lead to adjustments in goods prices in the local market, that is, there is no pass-through of exchange rate changes. Feenstra and Kendall (1997), Haskel and Wolf (2001), and Knetter (1993), among others, document the empirical relevance of LCP. Hence, it seems reasonable that PPP among traded goods should not simply be assumed,

[4]See Baum, Caglyan and Barkoulas (2001), Lo and Zivot (2001), Michael, Nobay and Peel (1997), Obstfeld and Taylor (1997), O'Connell and Wei (2002), Sarantis (1999), Taylor, Peel and Sarno (2001) and Taylor (2001). It should be noted that many of these papers consider the relative price of individual goods and therefore actually test the Law of One Price [LOP].

but rather should be tested. In this paper we show how this can be done with real exchange rates based on *aggregate* price indices, despite the fact these aggregate indices for a significant portion consist of non-traded goods, and therefore contain a nonstationary component. We show that this is no problem, and that the resultant test is simple, reliable and useful.

The paper proceeds as follows. In Section 2, we briefly summarize the consequences of the Balassa-Samuelson effect on real exchange rates based on price indices which include non-traded goods. In this section we also discuss how PPP among traded goods can be tested with such general real exchange rates. In Section 3, we apply the tests to real exchange rates and relative prices of several good categories with different degree of tradability across 13 developed countries. We find very little support for PPP, not even among traded goods. Finally, Section 4 concludes.

2 Testing for PPP while allowing for the Balassa-Samuelson effect

In this section we first briefly summarize the consequences of the intuitively plausible Balassa-Samuelson effect on testing for PPP among traded goods when general price indices are used to construct real exchange rates. Next, we discuss a simple statistical test method, and we show that is useful and reliable.

2.1 Incorporating the Balassa-Samuelson effect

Let q_t denote the logarithm of the real exchange rate, that is,

$$q_t = s_t + p_t^* - p_t, \tag{1}$$

where s_t is the logarithm of the nominal exchange rate, expressed in numbers of units of domestic currency needed to purchase one unit of foreign currency, and where p_t and p_t^* denote the logarithms of the domestic and foreign general price indices, respectively. These price indices are assumed to be geometric averages of traded and non-traded goods, that is,

$$p_t = (1-\alpha)p_t^T + \alpha p_t^N \qquad \text{and} \tag{2}$$

$$p_t^* = (1-\beta)p_t^{T*} + \beta p_t^{N*}, \tag{3}$$

where p_t^T and p_t^N denote the logarithms of the domestic price indices of traded and non-traded goods and α is the weight of non-traded goods in the general domestic price index p_t, while p_t^{T*}, p_t^{N*}, and β are defined similarly for the foreign country. Combining (2) and (3) with (1) allows the real exchange rate q_t to be decomposed as

$$q_t = x_t + y_t, \tag{4}$$

$$x_t = s_t + p_t^{T*} - p_t^T, \tag{5}$$

$$y_t = \beta(p_t^{N*} - p_t^{T*}) - \alpha(p_t^N - p_t^T). \tag{6}$$

In the Balassa-Samuelson model, permanent shocks to the relative productivity in traded and non-traded goods sectors lead to permanent changes in the relative prices of traded and non-traded goods. In present-day time series terminology, this implies that y_t is a non-stationary process. Consequently, the question of interest when testing for PPP is then whether the traded goods component x_t is stationary or not, or put differently, whether x_t contains a unit root or not. This can be represented formally, for example, by assuming that the stochastic processes for y_t and x_t are given by

$$y_t = y_{t-1} + \varepsilon_t, \tag{7}$$

$$x_t = \phi x_{t-1} + \eta_t, \qquad \text{with } -1 < \phi \leq 1, \tag{8}$$

where ε_t and η_t are i.i.d. random variables with mean zero and variances σ_ε^2 and σ_η^2, respectively, and with contemporaneous correlation $\rho = \sigma_{\varepsilon\eta}/(\sigma_\varepsilon\sigma_\eta)$. The null hypothesis of interest is in this case given by H_0: $\phi = 1$, which is to be tested against the alternative H_1: $\phi < 1$. In the sequel of this section we put forward a simple test for this hypothesis based on data for q_t only, where $q_t = x_t + y_t$.

2.2 A simple statistical test for PPP among traded goods

Proceeding from the above results, it is straightforward to show that (7) and (8) imply that $q_t = x_t + y_t$ can be represented as an ARMA(1,1) process,

$$\Delta q_t = \phi \Delta q_{t-1} + u_t + \theta u_{t-1}, \tag{9}$$

where Δ denotes the first differencing operator defined as $\Delta q_t \equiv q_t - q_{t-1}$ and $u_t \sim \mathsf{i.i.d}(0, \sigma_u^2)$. The moving average parameter θ is related to the parameters of the components x_t and y_t by[5]

$$\frac{\theta}{1+\theta^2} = \frac{-\phi\sigma_\varepsilon^2 - \sigma_\eta^2 - (1+\phi)\sigma_{\varepsilon\eta}}{(1+\phi^2)\sigma_\varepsilon^2 + 2\sigma_\eta^2 + 2(1+\phi)\sigma_{\varepsilon\eta}}. \tag{10}$$

[5]It is perhaps useful to mention that the corresponding expression for θ in Engel (2000) contains a misprint.

From (10), it is easy to see that under the null hypothesis $\mathsf{H}_0 : \phi = 1$, it holds that $\frac{\theta}{1+\theta^2} = -\frac{1}{2}$, which uniquely implies that $\theta = -1$. In that case, the autoregressive and moving average component in (9) cancel, and q_t reduces to a random walk

$$\Delta q_t = u_t. \tag{11}$$

Hence, the null hypothesis implies (11), while the alternative hypothesis implies (9). In sum, testing for PPP among traded goods in the presence of the Balassa-Samuelson effectively amounts to testing the null hypothesis of an ARMA(0,0) representation of the first differences of the real exchange rate q_t as in (11) against the alternative of an ARMA(1,1) representation as in (9).

Testing ARMA(0,0) against ARMA(1,1) is an intriguing statistical problem as there are so-called unidentified nuisance parameters under the null hypothesis. As discussed in general by Andrews and Ploberger (1994) and Hansen (1996), among others, conventional test statistics cannot be used in that case, and other strategies are required. Interestingly, Andrews and Ploberger (1996) developed statistical tests specifically designed for our testing problem. To describe their test statistics, consider the general ARMA(1,1) model for a time series z_t,

$$z_t = (\pi + \lambda)z_{t-1} + e_t - \pi e_{t-1}, \qquad t = 2, 3, \ldots, T, \tag{12}$$

where T is the sample size, and where e_t denotes a white noise process. The interest is in testing the null hypothesis of white noise against the alternative of ARMA(1,1)-type serial correlation in z_t, given by

$$\mathsf{H}_0 : \lambda = 0 \quad \text{and} \quad \mathsf{H}_1 : \lambda \neq 0,$$

respectively. The problem of unidentified nuisance parameters under the null hypothesis is immediately clear from the representation in (12), because in case $\lambda = 0$, the model reduces to $z_t = e_t$ and the parameter π has disappeared.

The testing approach devised by Andrews and Ploberger (1996) amounts to first computing a standard likelihood ratio (LR) or Lagrange Multiplier (LM) statistic for a large number of different *given* values of π, and then computing a certain functional of these "pointwise" statistics. Specifically, they consider the supremum, average and exponential statistics, given by

$$\text{SupF} = \sup_{\pi\in\Pi} F_T(\pi), \tag{13}$$

$$\text{AveF} = \frac{1}{n_\pi}\sum_{\pi\in\Pi} F_T(\pi), \tag{14}$$

$$\text{ExpF} = \ln\left(\frac{1}{n_\pi}\sum_{\pi\in\Pi}\exp\left(\frac{1}{2}F_T(\pi)\right)\right), \tag{15}$$

where F=LR or LM, where $F_T(\pi)$ denotes the standard LR or LM statistic of H_0: $\lambda = 0$ for a given π, and where Π is the set of all possible values of π and n_π is the number of selected elements of Π.[6] Andrews and Ploberger (1996) derive the non-standard asymptotic distributions of the supremum, average and exponential statistics and provide finite sample critical values.

In Monte Carlo simulation experiments reported by Andrews and Ploberger (1996), the SupLR and ExpLR tests are found to have very good power properties against various alternatives, followed by the ExpLM statistic. Here we perform an additional simulation experiment, to examine the properties of the test statistics when applied to a time series that is the sum of a stationary and a non-stationary component, as are the real exchange rates in the presence of the Balassa-Samuelson effect. Specifically, the DGP in the simulation experiment is given by (4), (7) and (8) and Δq_t plays the role of z_t in the tests. The autoregressive parameter for the stationary component x_t is varied among $\phi \in \{0.8, 0.85, 0.9, 0.95\}$. The standard deviation of the transitory shock η takes the values $\sigma_\eta \in \{1, 2, 4, 8, 16\}$, while the standard deviation of the permanent shock ε is set equal to 1 throughout. Both ε and η are assumed to be normally distributed. Finally, the correlation between ε_t and η_t is varied among $\rho \in \{0.0, 0.2, \ldots, 0.8\}$. All results are based on 1000 replications for sample size $T = 300$, which corresponds with the length of the real exchange rates considered in the empirical application below.

Following the recommendations of Andrews and Ploberger (1996), all tests are implemented using $\Pi = \{0, \pm.01, \pm.02, \ldots, \pm.79, \pm.80\}$ as the set of possible values of π. Furthermore, all series are demeaned. To obtain appropriate critical values, we simulated the finite sample distributions of the test statistics for this choice of Π using 100,000 replications of a driftless random walk. Table 1 shows rejection frequencies of the SupLR test at 10%, 5% and 1% significance levels[7].

[6] Explicit expressions for $LR_T(\pi)$ and $LM_T(\pi)$, which avoid the necessity of estimating the ARMA(1,1) model under the alternative, are also given in Andrews and Ploberger (1996).

[7] The corresponding critical values are 4.687, 6.039 and 9.245, respectively.

- insert Table 1 about here -

The results in Table 1 suggest that the power of the SupLR test not unexpectedly, decreases with ϕ and that it increases with ρ and σ_η. Note that the finding of Engel (1999) and Rogers and Jenkins (1995) that changes in relative prices of non-traded goods account for only a small fraction of real exchange rate changes, implies that in practice σ_η^2 is large relative to σ_ε^2. For example, Engel (2000) reports estimates based on quarterly US/UK data over the period 1970-1995 which imply that σ_η is 100 times larger than σ_ε^2, and under these conditions, our test should have considerable power. Finally, it seems to be recommendable to use a significance level of 10%, to retain at least a reasonable power for values of ϕ close to 1.

Detailed results for the remaining test statistics are not shown here to save space, but these are available upon request. In general, we find that the SupLR test has somewhat better power than the ExpLR test, which in turn performs slightly better than the AveLR test. The same relative performance is observed for the LM tests, and in each case their power is slightly less than the power of the corresponding LR statistics.

To conclude, in this section we have shown that PPP among traded goods can be tested with real exchange rates based upon aggregate price indices, using a simple test which seems to be reliable and useful.

3 Empirical Results

In this section we apply the Andrews-Ploberger test described above to relative price levels among a group of 13 industrialized countries[8]. The data are taken from the OECD *Main Economic Indicators*, and concern the general consumer price index (CPI), and the components of the CPI for (1) food, (2) all goods less food, (3) rent and (4) all services less rent. Note that these four CPI components are collectively exhaustive and mutually exclusive, and represent categories of goods with different degrees of tradability. The sample period covers the complete post-Bretton Woods era until the fixing of the Euro conversion rates, and runs from January 1973 until December 1998.[9] Monthly nominal exchange rate data are obtained from the IMF

[8]These countries and our mnemonics are: Belgium (BEL), Canada (CAN), Denmark (DNK), France (FRA), Germany (GER), Italy (ITA), Japan (JAP), the Netherlands (NLD), Norway (NOR), Spain (SPA), Switzerland (SWI), the United Kingdom (UK), and the United States (US).

[9]Exceptions are the UK all goods less food and all services less rent series, both of which end in July 1988; the Belgium rent series, which starts in June 1976; and the German all services less rent series, which starts in January 1991.

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As in the simulation experiment, all tests are computed on the demeaned changes in the real exchange rates, using $\Pi = \{0, \pm.01, \pm.02, \ldots, \pm.79, \pm.80\}$ as the set of possible values of π, while the appropriate finite sample distributions are used to assess the significance of the tests. Below we only report results for the SupLR test. Results for the remaining LR and LM statistics are very similar and are available upon request.

Table 2 contains results for all pairwise real exchange rates based upon the general CPIs. In this and all subsequent tables, the above-diagonal entries are values of the SupLR test statistic while the below-diagonal entries are the corresponding p-values based upon the simulated finite sample distribution.

- insert Table 2 about here -

At 1, 5 and 10% significance levels, the null hypothesis of no serial correlation can be rejected for 1, 10 and 15 real exchange rates, respectively, out of a total of 78 rates examined. Hence, it seems fair to draw the general conclusion that there is not much evidence in favor of PPP, although there are some exceptions. Notably for the UK, we find evidence (at the 10% level) of PPP with Belgium, Denmark, France, Germany, Japan and The Netherlands.

Tables 3-6 report results for real exchange rates based upon the CPI components for food, all goods less food, rent, and all services less rent.

- insert Tables 3-6 about here -

At the 10% significance level, there is evidence of ARMA(1,1)-type serial correlation in Δq_t, and hence of stationarity of the real exchange rate, in 14, 22, 19, and 17 of the 78 cases, respectively. This is well below the tabulated power levels of the test, and hence we can safely conclude that generally there is not much evidence of PPP among traded goods. Note that the number of rejections of the null hypotheses does not appear to be related to the degree of tradability of the goods and services in the different CPI components. For food, which might be regarded to have the smallest non-traded goods component, we actually find less evidence of PPP than for rent and services less rent. Again, there are some specific cases in favor of PPP, and in particular the results for the UK are interesting. Finally, comparing Tables 2 to 6, we notice that there are many non-overlapping cases with significant test statistics. This suggests that the choice for a price indicator can matter substantially.

4 Conclusions

In this paper we have put forward a simple test for PPP among traded goods, which could also be used for real exchange rates based on a general CPI. The test was shown to be reliable and useful in a simulation experiment with a DGP mimicking the composition of real exchange rates as the sum of a nonstationary and a stationary component. Upon application to bilateral real exchange rates among a group of 13 industrialized countries, we found little evidence in favor of PPP, although there were a few interesting exceptions, in particular for the UK.

As mentioned before, to increase power of tests for PPP, one can either increase the time span of the data or increase the cross-sectional dimension towards a panel of time series. We believe the latter strategy to be fruitful, also for our proposed testing strategy. Hence, a test for white noise against ARMA(1,1) for a panel of time series might provide even more powerful insights into any empirical evidence of PPP.

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Table 1: Power of SupLR test

		$\phi = 0.8$				$\phi = 0.85$				$\phi = 0.9$				$\phi = 0.95$			
σ_η	ρ	θ	10%	5%	1%	θ	10%	5%	1%	θ	10%	5%	1%	θ	10%	5%	1%
1	0.00	−0.86	0.46	0.33	0.14	−0.90	0.34	0.24	0.09	−0.93	0.24	0.14	0.04	−0.96	0.16	0.08	0.01
	0.20	−0.87	0.58	0.43	0.21	−0.90	0.41	0.29	0.11	−0.94	0.28	0.17	0.05	−0.97	0.17	0.09	0.02
	0.40	−0.88	0.64	0.52	0.27	−0.91	0.48	0.35	0.14	−0.94	0.30	0.20	0.06	−0.97	0.17	0.09	0.03
	0.60	−0.89	0.71	0.60	0.32	−0.92	0.52	0.38	0.17	−0.94	0.33	0.20	0.07	−0.97	0.19	0.10	0.03
	0.80	−0.89	0.77	0.64	0.36	−0.92	0.57	0.42	0.18	−0.95	0.34	0.22	0.08	−0.97	0.19	0.10	0.03
2	0.00	−0.95	0.90	0.80	0.55	−0.96	0.68	0.55	0.29	−0.98	0.42	0.29	0.10	−0.99	0.19	0.11	0.02
	0.20	−0.95	0.92	0.84	0.58	−0.97	0.71	0.59	0.32	−0.98	0.44	0.31	0.12	−0.99	0.20	0.11	0.01
	0.40	−0.96	0.93	0.85	0.63	−0.97	0.74	0.61	0.34	−0.98	0.47	0.33	0.13	−0.99	0.21	0.13	0.02
	0.60	−0.96	0.93	0.87	0.66	−0.97	0.77	0.63	0.38	−0.98	0.48	0.34	0.13	−0.99	0.21	0.13	0.03
	0.80	−0.96	0.94	0.89	0.68	−0.97	0.78	0.66	0.40	−0.98	0.48	0.35	0.14	−0.99	0.21	0.13	0.03
4	0.00	−0.99	1.00	0.96	0.83	−0.99	0.87	0.78	0.48	−0.99	0.54	0.40	0.17	−1.00	0.21	0.12	0.03
	0.20	−0.99	0.99	0.97	0.83	−0.99	0.87	0.77	0.48	−0.99	0.54	0.42	0.18	−1.00	0.22	0.13	0.03
	0.40	−0.99	0.99	0.97	0.84	−0.99	0.88	0.76	0.51	−0.99	0.56	0.42	0.19	−1.00	0.23	0.14	0.02
	0.60	−0.99	0.99	0.96	0.84	−0.99	0.87	0.77	0.52	−0.99	0.56	0.43	0.19	−1.00	0.24	0.14	0.03
	0.80	−0.99	0.98	0.96	0.84	−0.99	0.87	0.78	0.52	−0.99	0.56	0.42	0.18	−1.00	0.23	0.14	0.04
8	0.00	−1.00	1.00	0.98	0.90	−1.00	0.91	0.83	0.56	−1.00	0.59	0.44	0.20	−1.00	0.23	0.13	0.03
	0.20	−1.00	1.00	0.98	0.90	−1.00	0.91	0.83	0.55	−1.00	0.59	0.44	0.20	−1.00	0.23	0.12	0.03
	0.40	−1.00	1.00	0.99	0.90	−1.00	0.92	0.82	0.56	−1.00	0.59	0.45	0.20	−1.00	0.24	0.13	0.03
	0.60	−1.00	1.00	0.98	0.88	−1.00	0.91	0.82	0.56	−1.00	0.59	0.47	0.21	−1.00	0.24	0.15	0.03
	0.80	−1.00	0.99	0.97	0.88	−1.00	0.90	0.81	0.57	−1.00	0.59	0.46	0.20	−1.00	0.24	0.15	0.05
16	0.00	−1.00	1.00	0.99	0.91	−1.00	0.92	0.85	0.58	−1.00	0.60	0.45	0.21	−1.00	0.23	0.12	0.03
	0.20	−1.00	1.00	0.99	0.90	−1.00	0.92	0.85	0.58	−1.00	0.60	0.45	0.21	−1.00	0.23	0.13	0.03
	0.40	−1.00	1.00	0.99	0.91	−1.00	0.93	0.83	0.58	−1.00	0.60	0.46	0.21	−1.00	0.24	0.13	0.03
	0.60	−1.00	1.00	0.98	0.90	−1.00	0.92	0.83	0.58	−1.00	0.60	0.46	0.21	−1.00	0.24	0.15	0.03
	0.80	−1.00	1.00	0.98	0.89	−1.00	0.90	0.83	0.59	−1.00	0.61	0.47	0.22	−1.00	0.25	0.15	0.05

Notes: The table contains rejection frequencies of the SupLR statistic (13) with $\Pi = \{0, \pm.01, \pm.02, \ldots, \pm.79, \pm.80\}$ at the 5% significance level when applied to the first differences of q_t generated according to (4), (7) and (8), with ε_t and η_t normally distributed and $\sigma_\varepsilon^2 = 1$. The table is based on 1000 replications for sample size $T = 300$.

Table 2: Andrews-Ploberger SupLR test for pairwise PPP using real exchange rates based on general CPI

	BEL	CAN	DNK	FRA	GER	ITA	JPN	NLD	NOR	SPA	SWI	UK	US
BEL		2.79	5.28	5.69	12.00	2.73	3.37	1.60	3.24	1.43	2.38	8.27	1.27
CAN	0.26		2.28	2.38	1.91	3.17	3.07	1.59	0.36	2.63	0.85	2.11	0.55
DNK	0.07	0.34		0.98	2.29	2.40	2.91	7.61	1.44	4.00	2.05	4.93	1.04
FRS	0.06	0.33	0.68		3.60	1.07	4.19	4.40	1.71	4.88	1.95	7.74	0.78
GER	0.00	0.42	0.34	0.17		3.33	2.67	8.60	5.34	1.11	3.24	8.70	0.81
ITA	0.27	0.22	0.32	0.65	0.20		6.56	1.09	0.21	3.61	1.95	3.38	2.43
JPN	0.20	0.23	0.25	0.13	0.28	0.04		2.10	2.66	6.34	1.49	8.60	3.67
NLD	0.49	0.49	0.02	0.12	0.01	0.64	0.38		3.86	2.53	1.83	8.60	0.48
NOR	0.21	0.91	0.53	0.47	0.07	0.96	0.28	0.15		3.30	1.16	0.53	0.25
SPA	0.54	0.29	0.14	0.09	0.63	0.17	0.04	0.30	0.20		1.48	1.42	2.59
SWI	0.33	0.72	0.39	0.41	0.21	0.41	0.52	0.44	0.62	0.52		2.95	1.58
UK	0.02	0.38	0.09	0.02	0.01	0.19	0.01	0.01	0.84	0.54	0.24		2.27
US	0.58	0.83	0.66	0.75	0.74	0.32	0.17	0.86	0.95	0.29	0.50	0.35	

Notes: The above-diagonal entries are values of the Andrews-Ploberger SupLR test statistic applied to real exchange rates based on general CPI's, using monthly data for the period 1973.1-1998.12. The below-diagonal entries are the corresponding p-values.

Table 3: Andrews-Ploberger SupLR test on relative prices of food

	BEL	CAN	DNK	FRA	GER	ITA	JPN	NLD	NOR	SPA	SWI	UK	US
BEL		1.87	4.07	10.24	9.72	2.21	3.75	0.98	0.59	1.58	2.23	6.97	0.80
CAN	0.43		1.45	0.52	0.94	1.98	2.62	0.57	0.82	0.58	0.45	1.48	4.27
DNK	0.14	0.53		1.11	2.79	2.38	2.96	8.22	1.55	3.89	2.17	5.24	1.32
FRA	0.01	0.85	0.63		3.50	0.75	2.18	2.14	3.52	7.47	4.83	4.53	0.19
GER	0.01	0.69	0.26	0.18		3.71	3.35	3.14	6.21	1.61	8.09	4.06	0.65
ITA	0.36	0.40	0.33	0.76	0.16		6.84	1.00	1.45	6.81	1.38	1.85	2.41
JPN	0.16	0.29	0.24	0.36	0.20	0.03		1.17	1.43	4.77	2.19	8.46	5.25
NLD	0.68	0.83	0.02	0.37	0.22	0.67	0.61		1.50	5.37	3.61	3.62	0.18
NOR	0.82	0.73	0.50	0.18	0.05	0.53	0.54	0.52		3.24	2.30	2.27	0.83
SPA	0.50	0.82	0.15	0.02	0.49	0.03	0.10	0.07	0.21		2.10	1.55	0.91
SWI	0.35	0.88	0.36	0.09	0.02	0.55	0.36	0.17	0.34	0.38		1.38	0.92
UK	0.03	0.52	0.08	0.11	0.14	0.43	0.02	0.17	0.35	0.50	0.55		3.57
US	0.74	0.12	0.57	0.97	0.80	0.32	0.08	0.97	0.73	0.70	0.70	0.18	

Notes: The above-diagonal entries are values of the Andrews-Ploberger SupLR test statistic applied to the relative prices of food expressed in a common currency, using monthly data for the period 1973.1-1998.12. The below-diagonal entries are the corresponding p-values.

Table 4: Andrews-Ploberger SupLR test on relative prices of all goods less food

	BEL	CAN	DNK	FRA	GER	ITA	JPN	NLD	NOR	SPA	SWI	UK	US
BEL		2.21	24.97	1.97	11.64	1.92	2.58	9.41	12.93	1.10	1.08	6.59	0.65
CAN	0.36		1.53	2.54	2.45	4.26	2.67	1.04	0.30	3.73	0.96	1.48	1.75
DNK	0.00	0.51		2.79	10.35	1.40	1.69	21.35	8.20	4.44	6.79	2.76	0.72
FRA	0.41	0.30	0.26		3.83	1.43	4.06	4.41	8.17	2.10	1.83	7.74	0.86
GER	0.00	0.31	0.01	0.15		3.87	2.56	13.70	11.30	0.17	3.71	9.32	1.27
ITA	0.42	0.12	0.55	0.54	0.15		6.43	3.06	1.72	1.43	2.08	5.32	3.13
JPN	0.29	0.28	0.47	0.14	0.30	0.04		1.35	1.89	5.90	1.05	7.34	3.76
NLD	0.01	0.66	0.00	0.11	0.00	0.23	0.56		18.85	1.82	5.42	8.66	0.34
NOR	0.00	0.93	0.02	0.02	0.00	0.46	0.42	0.00		7.15	1.00	1.07	0.99
SPA	0.64	0.16	0.11	0.38	0.97	0.54	0.05	0.44	0.03		0.94	1.61	2.74
SWI	0.64	0.68	0.03	0.44	0.16	0.38	0.66	0.07	0.67	0.69		1.95	1.75
UK	0.04	0.52	0.27	0.02	0.01	0.07	0.03	0.01	0.65	0.49	0.41		1.57
US	0.80	0.46	0.77	0.72	0.58	0.22	0.16	0.91	0.67	0.27	0.45	0.50	

Notes: The above-diagonal entries are values of the Andrews-Ploberger SupLR test statistic applied to the relative prices of all goods less food expressed in a common currency, using monthly data for the period 1973.1-1998.12. The below-diagonal entries are the corresponding p-values.

Table 5: Andrews-Ploberger SupLR test on relative prices of rent

	BEL	CAN	DNK	FRA	GER	ITA	JPN	NLD	NOR	SPA	SWI	UK	US
BEL		3.57	12.2	0.42	1.10	0.63	1.60	2.08	1.13	0.82	2.86	22.88	1.76
CAN	0.18		0.97	2.28	2.31	4.93	3.97	0.54	0.24	3.84	0.60	4.76	1.16
DNK	0.00	0.68		10.43	15.03	2.66	1.24	28.53	9.58	3.99	0.99	2.49	1.20
FRA	0.88	0.35	0.01		1.29	2.01	2.85	5.12	1.03	0.52	1.78	14.10	0.91
GER	0.64	0.34	0.00	0.58		1.93	1.95	19.29	1.70	0.56	0.72	17.68	0.89
ITA	0.81	0.09	0.28	0.40	0.41		5.90	0.07	0.75	0.96	1.09	8.13	4.30
JPN	0.49	0.14	0.59	0.26	0.41	0.05		0.32	2.70	9.92	0.33	13.73	3.10
NLD	0.38	0.84	0.00	0.08	0.00	0.99	0.92		3.23	1.30	7.74	13.69	0.44
NOR	0.63	0.95	0.01	0.66	0.47	0.76	0.28	0.21		1.85	8.38	0.83	1.53
SPA	0.73	0.15	0.14	0.85	0.83	0.68	0.01	0.58	0.43		0.32	2.47	3.41
SWI	0.25	0.82	0.68	0.45	0.77	0.64	0.92	0.02	0.02	0.92		4.32	1.67
UK	0.00	0.10	0.31	0.00	0.00	0.02	0.00	0.00	0.73	0.31	0.12		4.34
US	0.45	0.62	0.61	0.70	0.71	0.12	0.23	0.88	0.51	0.19	0.47	0.12	

Notes: The above-diagonal entries are values of the Andrews-Ploberger SupLR test statistic applied to the relative prices of rent expressed in a common currency, using monthly data for the period 1973.1-1998.12. The below-diagonal entries are the corresponding p-values.

Table 6: Andrews-Ploberger SupLR test on relative prices of all services less rent

	BEL	CAN	DNK	FRA	GER	ITA	JPN	NLD	NOR	SPA	SWI	UK	US
BEL		6.12	1.49	5.62	0.86	4.60	1.86	0.42	5.65	0.30	0.95	6.98	3.77
CAN	0.05		3.11	4.87	1.84	3.48	4.03	2.88	1.42	3.21	1.84	5.19	1.14
DNK	0.52	0.22		0.67	3.30	5.28	1.23	2.88	0.56	2.86	4.54	5.06	2.05
FRA	0.06	0.09	0.79		2.23	2.67	2.88	2.66	5.04	1.39	1.32	5.27	2.17
GER	0.72	0.43	0.20	0.35		1.96	0.23	14.42	0.37	5.08	2.34		5.37
ITA	0.10	0.18	0.07	0.28	0.41		4.29	2.08	1.45	3.02	2.29	1.68	3.35
JPN	0.43	0.14	0.60	0.25	0.96	0.12		1.40	1.53	6.61	1.94	4.98	3.02
NLD	0.89	0.25	0.25	0.28	0.00	0.38	0.55		2.51	1.31	1.38	5.09	1.53
NOR	0.06	0.54	0.83	0.08	0.90	0.53	0.51	0.31		2.50	0.43	2.27	1.58
SPA	0.93	0.21	0.25	0.55	0.08	0.23	0.04	0.57	0.31		0.90	2.82	3.29
SWI	0.69	0.43	0.11	0.57	0.33	0.34	0.41	0.55	0.88	0.71		2.81	1.92
UK	0.03	0.08	0.08	0.07		0.47	0.09	0.08	0.35	0.26	0.26		1.17
US	0.16	0.62	0.39	0.36	0.07	0.20	0.23	0.51	0.50	0.20	0.42	0.62	

Notes: The above-diagonal entries are values of the Andrews-Ploberger SupLR test statistic applied to the relative prices of all services less rent expressed in a common currency, using monthly data for the period 1973.1-1998.12. The below-diagonal entries are the corresponding p-values.