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Long-horizon regression tests of the theory of purchasing power parity

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Abstract

In this article we test the purchasing power parity (PPP) hypothesis during the recent floating exchange rate period, using quarterly data for 21 OECD countries. In doing so, we use the long-horizon regression approach developed by Fisher and Seater [American Economic Review 83 (1993) 402] and consider 60 bilateral intercountry relations. We investigate the power of the long-horizon regression tests, using the inverse power function of Andrews [Econometrica 57 (1989) 1059], and provide weak evidence in favor of PPP.

JEL classification: C22; F31
Keywords: Purchasing power parity; Integration; Long-run derivative; Inverse power function

1. Introduction

The theory of purchasing power parity (PPP) has attracted a great deal of attention and has been explored extensively in the recent literature using recent advances in the field of applied econometrics (that pay explicit attention to the integration and cointegration properties of the variables). Based on the law of one price, PPP asserts that relative goods prices are not affected by exchange rates – or, equivalently, that exchange rate changes will be proportional to relative inflation. The relationship is important not only because it has been a cornerstone of exchange rate models in

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international economics, but also because of its policy implications; it provides a benchmark exchange rate and hence has some practical appeal for policymakers and exchange rate arbitragers.

Empirical studies generally fail to find support for purchasing power parity, especially during the recent floating exchange rate period. In fact, the empirical consensus is that PPP does not hold over this period (see, for example, Adler and Lehman, 1983; Mark, 1990; Patel, 1990; Grilli and Kaminsky, 1991; Flynn and Boucher, 1993; Serletis, 1994; Serletis and Zimonopoulos, 1997; Coe and Serletis, 2002). But there are also studies covering different groups of countries as well as studies covering periods of long duration or country pairs experiencing large differentials in price movements that report evidence of mean reversion towards PPP (see, for example, Frenkel, 1980; Diebold et al., 1991; Glen, 1992; Perron and Vogelsang, 1992; Phylaktis and Kassimatis, 1994; Lothian and Taylor, 1996). Also, studies using high-frequency (monthly) data over the recent floating exchange rate period report significant evidence favorable to purchasing power parity (see, for example, Pippenger, 1993; Cheung and Lai, 1993; Kugler and Lenz, 1993). [1]

Although purchasing power parity has been studied extensively, recently Fisher and Seater (1993) contribute to the literature on testing key classical macroeconomic hypotheses (such as, for example, the neutrality of money proposition, the Fisher relation, and a vertical long-run Phillips curve) by developing tests (using recent advances in the theory of nonstationary regressors) based on coefficient restrictions in bivariate vector autoregressive models. They show that meaningful tests can only be constructed if the relevant variables satisfy certain nonstationarity conditions and that much of the older literature violates these requirements, and hence has to be disregarded.

In this paper we adopt the long-horizon regression approach of Fisher and Seater (1993) for studying the purchasing power parity proposition. Long-horizon regressions have received a lot of attention in the recent economics and finance literature, because studies based on long-horizon variables seem to find significant results where short-horizon regressions commonly used in economics and finance have failed.

We use quarterly data, over the period from 1973:1 to 1998:4, for 21 OECD countries and pay particular attention to the integration properties of the variables, since meaningful purchasing power parity tests require that both the nominal exchange rate and the relative price satisfy certain nonstationarity conditions. The countries involved are Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Japan, the Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, the United Kingdom, and the United States.

Our analysis is organized as follows. Section 2 presents a brief summary of the purchasing power parity hypothesis and reviews earlier empirical tests of the hypothesis. Section 3 provides a summary of the long-horizon regression approach devel-

[1] There are also recent studies that use panel methods, such as, for example, Koedijk et al. (1998) and Papell and Theodoridis (1998), as well as studies that consider the effect of transaction costs and nonlinear adjustments, such as, for example, Michael et al. (1997), that report evidence favorable to PPP.

oped by Fisher and Seater (1993). Section 4 discusses the data and presents the long-horizon regression test results with inference based on 95% confidence bands, constructed using the Newey and West (1987) procedure. Section 5 investigates the power of the long-horizon regression tests using the Andrews (1989) inverse power function. Section 6 investigates the robustness of our results to the use of alternative price indices and the final section closes with a brief summary and conclusion.

2. Earlier tests of PPP

Consider the following relationship between the (equilibrium) domestic currency value of one unit of foreign currency and the domestic and foreign price levels:

$$S_t = A\frac{P_t}{P_t^*}, \tag{1}$$

where A is an arbitrary constant term, S_t denotes the nominal exchange rate (domestic currency value per unit of foreign currency), P_t the domestic price level (in domestic currency), and P_t^* the foreign price level (in foreign currency). Taking logarithms we arrive at the linear relationship

$$s_t = \alpha + p_t - p_t^*, \tag{2}$$

where α, s, p, and p^* denote logarithms of A, S, P, and P^*, respectively. The 'absolute version' of PPP states that $A = 1$, or equivalently, that (2) holds with the additional restriction $\alpha = 0$. [2]

A natural test of absolute purchasing power parity – see Frenkel (1980) – is to estimate the equation

$$s_t = \alpha + \beta p_t - \beta^* p_t^* + u_t \tag{3}$$

and test the hypothesis that the constant term is zero, i.e., $\alpha = 0$ and the hypothesis that the coefficients of domestic and foreign prices are both unity (as implied by Eq. (2)), i.e., $\beta = \beta^* = 1$. [3]

[2] The 'relative version' of PPP, relating the change in the exchange rate to the inflation rates in the two countries, can be written as

$$\Delta s_t = \Delta p_t - \Delta p_t^*,$$

where Δp_t and Δp_t^* denote the domestic and foreign inflation rate, respectively – see Rogoff (1996) for more details.

[3] In its relative price formulation, purchasing power parity can be tested by estimating the equation

$$\Delta s_t = a + b\Delta p_t - b^*\Delta p_t^* + \xi_t$$

and testing the hypothesis that the constant term is zero, i.e., $a = 0$, as well as the hypothesis that the coefficients of domestic and foreign inflation rates are both unity, i.e., $b = b^* = 1$.

2.1. PPP as a cointegrating relation

Since PPP is a (backward looking) long-run relation, it can also be formulated as a testable hypothesis in terms of cointegration. Let Y_t be a multivariate stochastic process consisting of the logarithm of the nominal exchange rate (s_t), the domestic price level (p_t) and the foreign price level (p_t^*). Assuming that each component of Y_t is integrated of order one (or $I(1)$ in the terminology of Engle and Granger (1987)), the absolute purchasing power parity theory implies that the PPP relation is stationary. That is

$$\beta' Y_t = \begin{bmatrix} 1 & -1 & 1 \end{bmatrix} \begin{bmatrix} s_t \\ p_t \\ p_t^* \end{bmatrix} = \text{stationary}. \tag{4}$$

In the terminology of Engle and Granger (1987), the row vector β' is called cointegrating vector for the nonstationary stochastic process Y_t. This cointegrating vector isolates (in the present context) stationary linear combinations of the nonstationary stochastic process Y_t corresponding to s_t, p_t, and p_t^*.

There are different approaches available to test the hypothesis that the PPP relation is stationary. First, one may directly test the stationarity of $\beta' Y_t$ by univariate unit root tests. This approach, however, ignores the dynamic interrelationship of s_t, p_t, and p_t^*. Second, one may test for cointegration between s_t, p_t, and p_t^* using the Engle and Granger (1987) methodology – this involves selecting arbitrarily a normalization and regressing one variable on the others to obtain the OLS regression residuals and testing for a unit root in these residuals. This approach, however, does not distinguish between the existence of one or more cointegrating vectors, and the OLS parameter estimates of the cointegrating vector depend on the arbitrary normalization implicit in the selection of the dependent variable in the regression equation. Moreover, this approach breaks down when p_t and p_t^* cointegrate. These problems can be avoided by using Johansen's (1988) maximum likelihood extension of the Engle and Granger (1987) cointegration approach as, for example, in Johansen and Juselius (1992); Cheung and Lai (1993); Kugler and Lenz (1993), and Serletis (1994).

However, both the Engle and Granger (1987) and Johansen (1988) approaches are two-stage testing procedures of long-run absolute PPP. In the first stage, the null hypothesis of no cointegration is tested against the alternative of cointegration with an unknown cointegrating vector. If the null is rejected, a second-stage test is carried out with cointegration maintained under both the null and alternative. In particular, in the second stage, the null is that the data cointegrate with the specific cointegrating vector implied by long-run PPP (that is, $[1 \quad -1 \quad 1]$ in the case of the trivariate system, s_t, p_t, and p_t^*, or $[1 \quad -1]$ in the case of the bivariate system s_t and $p_t - p_t^*$) and the alternative is that the data cointegrate with another unspecified cointegrating vector.

2.2. PPP and the real exchange rate

Instead of testing for cointegration, a different approach to test the hypothesis that the PPP relation is stationary would be to compute a linear combination of the PPP theory variables (such as, for example, the real exchange rate) and investigate its univariate time series properties using usual unit root testing procedures. Such a test (directly on the real exchange rate) is actually a test for cointegration between nominal prices and the nominal exchange rate, but with a common factor restriction on the individual underlying dynamics and nominal exchange rate. [4]

The real exchange rate, E_t, can be calculated as

$$E_t = S_t \frac{P_t^*}{P_t}, \tag{5}$$

with S_t given by Eq. (1). Taking logarithms of Eq. (5), the real exchange rate becomes a linear combination of the nominal exchange rate and the domestic and foreign price levels, as follows:

$$e_t = s_t + p_t^* - p_t,$$

where e_t is the logarithm of E_t and s_t is given by Eq. (2). Clearly, under long-run purchasing power parity, the long-run equilibrium real exchange rate is equal to 1 (at every point in time) in the absolute version of PPP, which would imply $e_t = 0$. [5] In the short-run, however, we expect deviations from PPP, coming from stochastic shocks, and the question at issue is whether these deviations are permanent or transitory.

A sufficient condition for a violation of absolute PPP is that the real exchange rate is characterized by a unit root. A number of approaches have been developed to test for unit roots. Nelson and Plosser (1982), using augmented Dickey–Fuller (ADF) type regressions (see Dickey and Fuller, 1981), argue that most macroeconomic time series (including real exchange rates) have a unit root. Perron (1989), however, has shown that conventional unit root tests are biased against rejecting a unit root where there is a break in a trend stationary process. Motivated by these considerations, Serletis and Zimonopoulos (1997), using the methodology suggested by Perron and Vogelsang (1992) and quarterly dollar-based and Deutschemark-based real exchange rates (over the period from 1957:1 to 1995:4) for 17 OECD countries, show that the unit root hypothesis cannot be rejected even if allowance is made for the possibility of a one-time change in the mean of the series at an unknown point in time.

However, the (apparent) random walk behaviour of the real exchange rate could be contrasted with chaotic dynamics. This is motivated by the notion that the real exchange rate follows a deterministic nonlinear process which generates output that

[4] Such restrictions on the dynamics, however, could bias the results towards the null of a unit root (see, for example, Kremers et al., 1992).

[5] In the relative version of PPP the first logged difference of the real exchange rate would be zero, that is $\Delta e_t = 0$.

mimics the output of stochastic systems. In other words, it is possible for the real exchange rate to appear to be random but not to be really random. In fact, Serletis and Gogas (2000) test for chaos, using the Nychka et al. (1992) test (for positivity of the dominant Lyapunov exponent), in the dollar-based real exchange rate series used by Serletis and Zimonopoulos (1997), and find evidence of nonlinear chaotic dynamics in 7 out of 15 real exchange rate series. This suggests that real exchange rate movements might not be really random and that it is perhaps possible to model (by means of differential/difference equations) the nonlinear chaos generating mechanism and build a predictive model of real exchange rates – see Barnett and Serletis (2000) for some thoughts along these lines. [6]

3. The long-horizon regression approach

In this paper we test the theory of PPP using the long-horizon regression approach developed by Fisher and Seater (1993). One important advantage to working with the long-horizon regression approach is that cointegration is neither necessary nor sufficient for tests on the long-run derivative.

We start with the following bivariate autoregressive representation:

$$\alpha_{ss}(L)\Delta^{\langle s\rangle}s_t = \alpha_{sx}(L)\Delta^{\langle x\rangle}x_t + \varepsilon_t^s,$$
$$\alpha_{xx}(L)\Delta^{\langle x\rangle}x_t = \alpha_{xs}(L)\Delta^{\langle s\rangle}s_t + \varepsilon_t^x,$$

where $\alpha_{ss}^0 = \alpha_{xx}^0 = 1$, $\Delta = 1 - L$, where L is the lag operator, s_t is the nominal exchange rate, x_t is the relative price, $p_t - p_t^*$, and $\langle z\rangle$ represents the order of integration of z, so that if z is integrated of order γ (or $I(\gamma)$ in the terminology of Engle and Granger (1987)), then $\langle z\rangle = \gamma$ and $\langle \Delta z\rangle = \langle z\rangle - 1$. The vector $(\varepsilon_t^s, \varepsilon_t^x)'$ is assumed to be independently and identically distributed normal with zero mean and covariance $\sum_\varepsilon$, the elements of which are $\text{var}(\varepsilon_t^s)$, $\text{var}(\varepsilon_t^x)$, $\text{cov}(\varepsilon_t^s, \varepsilon_t^x)$.

According to this approach, purchasing power parity can be tested in terms of the long-run derivative of s_t with respect to a permanent change in x_t, which is defined as follows. If $\lim_{k\to\infty} \partial x_{t+k}/\partial \varepsilon_t^x \neq 0$, then

$$\text{LRD}_{s,x} = \lim_{k\to\infty} \frac{\partial s_{t+k}/\partial \varepsilon_t^x}{\partial x_{t+k}/\partial \varepsilon_t^x}.$$

Thus, in the present context $\text{LRD}_{s,x}$ expresses the ultimate effect of an exogenous relative price disturbance on the nominal exchange rate, s_t, relative to that distur-

[6] We also explored the presence of nonlinearities in the real exchange rate using a nonlinear, two-regime, self-exciting threshold autoregressive (SETAR) model for the real exchange rate. Following Potter (1995) and Hansen (1996), we estimated the model using least squares, allowing the threshold parameter to vary from the 15th to the 85th percentile of the empirical distribution of Δe_t and the delay parameter from 1 to 5. We tested the no threshold effect (single regime) null hypothesis, using the six LM-based tests used by Hansen (1996). For all 60 U.S. dollar-based, DM-based, and Japanese yen-based real exchange rate series (used in this paper) we reject the linear null model at the 1% level (these results are available from the authors upon request).

bance's ultimate effect on the relative price, x_t. When $\lim_{k\to\infty} \partial x_{t+k}/\partial\varepsilon_t^x = 0$, there are no permanent changes in x_t (i.e., x_t is $I(0)$) and thus $\mathrm{LRD}_{s,x}$ is undefined. In terms of this framework, purchasing power parity requires that $\mathrm{LRD}_{s,x} = 1$.

The above bivariate autoregressive system can be inverted to yield the following vector moving average representation:

$$\Delta^{\langle s\rangle} s_t = \theta_{sx}(L)\varepsilon_t^x + \theta_{ss}(L)\varepsilon_t^s,$$
$$\Delta^{\langle x\rangle} x_t = \theta_{xx}(L)\varepsilon_t^x + \theta_{xs}(L)\varepsilon_t^s.$$

In terms of this moving average representation, Fisher and Seater (1993) show that $\mathrm{LRD}_{s,x}$ depends on $\langle x\rangle - \langle s\rangle$, as follows:

$$\mathrm{LRD}_{s,x} = \frac{(1-L)^{\langle x\rangle - \langle s\rangle}\theta_{sx}(L)|_{L=1}}{\theta_{xx}(1)}.$$

Hence, meaningful purchasing power parity tests can be conducted if both s_t and x_t satisfy certain nonstationarity conditions. In particular, purchasing power parity tests require that both s_t and x_t are at least $I(1)$ and of the same order of integration. In fact, when $\langle s\rangle = \langle x\rangle = 1$, the long-run derivative becomes

$$\mathrm{LRD}_{s,x} = \frac{\theta_{sx}(1)}{\theta_{xx}(1)},$$

where $\theta_{sx}(1) = \sum_{j=1}^{\infty}\theta_{sx}^j$ and $\theta_{xx}(1) = \sum_{j=1}^{\infty}\theta_{xx}^j$. Above, the coefficient $\theta_{sx}(1)/\theta_{xx}(1)$ is the long-run value of the impulse-response of s_t with respect to x_t, implying that $\mathrm{LRD}_{s,x}$ is equivalent to the long-run elasticity of s_t with respect to x_t.

Under the assumptions that $\mathrm{cov}(\varepsilon_t^s, \varepsilon_t^x) = 0$ and that the relative price is exogenous in the long-run, the coefficient $\theta_{sx}(1)/\theta_{xx}(1)$ equals the zero-frequency regression coefficient in the regression of $\Delta^{\langle s\rangle}s$ on $\Delta^{\langle x\rangle}x$ – see Fisher and Seater (1993, note 11). This estimator is given by $\lim_{k\to\infty} b_k$, where b_k is the coefficient from the regression

$$\left[\sum_{j=0}^{k}\Delta^{\langle s\rangle}s_{t-j}\right] = a_k + b_k\left[\sum_{j=0}^{k}\Delta^{\langle x\rangle}x_{t-j}\right] + e_{kt}.$$

In fact, when $\langle s\rangle = \langle x\rangle = 1$, consistent estimates of b_k can be derived by applying ordinary least squares to the regression of the growth rate of s on the growth rate of x,

$$s_t - s_{t-k-1} = a_k + b_k[x_t - x_{t-k-1}] + e_{kt}, \quad k = 1,\ldots,K. \tag{6}$$

The null of purchasing power parity is $b_k = 1$. If the null is not rejected across a range of k-forecast horizons, the data supports purchasing power parity.

In the context of Eq. (6), one could also reverse the hypothesis and test the null that there is no PPP. However, the null of no PPP is $b_k = \phi$, $\forall\phi \neq 1$, and there are infinitely many possible null hypotheses to be tested. For this reason, we test the null that there is PPP, although the test might break down in some cases where we expect the null hypothesis to hold.

4. Data and the LRD tests

The data, taken from the IMF International Financial Statistics, consist of quarterly nominal exchange rates and consumer price indices covering the period 1973:1 to 1998:4 for 21 OECD countries. A final demand price is used in the calculation of the relative price instead of an output price, because of unavailability of the same quarterly output price for each country. It is to be noticed, however, that Perron and Vogelsang (1992) argue that the results for PPP tests could depend on the price used.

The countries involved are Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Japan, the Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, the United Kingdom, and the United States. In investigating purchasing power parity, however, we consider 60 bilateral intercountry relations; twenty relations between the United States as the home country and the other countries as the foreign countries; 20 relations between Germany as the home country and the other countries as the foreign countries; and 20 relations between Japan as the home country and the other countries as the foreign countries.

As was argued in the previous section, meaningful long-horizon regression tests of purchasing power parity require that both s_t and x_t are at least integrated of order one and of the same order of integration. We investigate this issue by calculating p-values [based on the response surface estimates given by MacKinnon (1994)] for the augmented Dickey–Fuller (ADF) test – see Dickey and Fuller (1981) and report the results in panel A of Tables 1–3. [7] Based on these p-values, the null hypothesis of a unit root in log levels cannot be rejected at the 5% level, except for 10 out of the 120 series being tested.

In particular, we reject the null hypothesis of a unit root with the DM-based exchange rates for Austria and the Netherlands and the yen-based exchange rates for Canada, Finland, and Switzerland. In each of these cases, $\mathrm{LRD}_{s,x} = 0$ by definition, since there are no permanent changes in the exchange rate and purchasing power parity is violated. Moreover, we reject the null hypothesis of a unit root with the U.S.-based relative prices for Belgium, Japan, and the United Kingdom, the Germany-based relative price for the Netherlands, and the Japan-based relative price for the United States. In each of these cases, $\mathrm{LRD}_{s,x}$ cannot be defined. It is to be noted that traditional explanations of purchasing power parity would favor acceptance of purchasing power parity for the bilateral Germany–Netherlands relation, at least more so than for many of the other bilateral relations we analyze in this paper. In Table 2, however, the null hypotheses of a unit root in the DM-based exchange rate for the Netherlands as well as the Germany-based relative price for the Netherlands are rejected at the 5% level. Although just a single case, the fact that the

[7] The optimal lag length was taken to be the order selected by the Akaike Information Criterion (AIC) plus 2 – see Pantula et al. (1994) for details regarding the advantages of this rule for choosing the number of augmenting lags.

Table 1
Unit root and cointegration tests with the United States as the home country

Country	(A) Unit root tests		(B) Cointegration tests, dependent variable	
	s_t	x_t	s_t	x_t
Australia	0.551	0.878	0.735	0.367
Austria	0.378	0.621	0.497	0.779
Belgium	0.507	0.000	–	–
Canada	0.517	0.889	0.709	0.925
Denmark	0.589	0.858	0.517	0.828
Finland	0.271	0.961	0.308	0.919
France	0.735	0.773	0.598	0.897
Germany	0.357	0.903	0.495	0.944
Greece	0.845	0.891	0.760	0.922
Ireland	0.457	0.768	0.304	0.897
Italy	0.639	0.835	0.612	0.924
Japan	0.287	0.000	–	–
Netherlands	0.426	0.237	0.608	0.366
New Zealand	0.492	0.985	0.228	0.655
Norway	0.422	0.840	0.356	0.811
Portugal	0.897	0.992	0.926	0.808
Spain	0.615	0.709	0.600	0.712
Sweden	0.512	0.972	0.197	0.852
Switzerland	0.142	0.460	0.307	0.787
United Kingdom	0.135	0.000	–	–

Note: Numbers are tail areas of tests.

unit root null is rejected for both unit root tests is not surprising, but our model breaks down for this bilateral relation, since $\mathrm{LRD}_{s,x}$ is undefined.

Although cointegration is neither necessary nor sufficient for tests on the long-run derivative, we also test the null hypothesis of no cointegration (against the alternative of cointegration) between s_t and x_t for those countries for which both s_t and x_t are integrated of order one, since cointegration is a property of integrated variables. In doing so, we use the Engle and Granger (1987) two-step procedure. The tests are first done with the nominal exchange rate s_t as the dependent variable in the cointegrating regression and then repeated with the relative price x_t as the dependent variable, with the number of augmenting lags chosen using the AIC + 2 rule mentioned earlier. The asymptotic p-values, using the coefficients in MacKinnon (1994), are reported in panel B of Tables 1–3 and suggest that the null hypothesis of no cointegration between s_t and x_t cannot be rejected (at the 5% level), except for France when Germany is used as the home country.

We start by estimating Eq. (6) for each of the twenty one countries with the United States, Germany, and Japan as home countries. We report results only for those countries for which $\mathrm{LRD}_{s,x}$ can be defined. As in Fisher and Seater (1993), we consider values of k ranging from 1 to 30 and present the estimates of b_k along with the 95% confidence bands in Figs. 1–21. The confidence bands are constructed using the Newey and West (1987) procedure from a t-distribution with T/k degrees of freedom, where T is the number of observations. In each of the 21 figures, in panel A

Table 2
Unit root and cointegration tests with Germany as the home country

Country	(A) Unit root tests		(B) Cointegration tests, dependent variable	
	s_t	x_t	s_t	x_t
Australia	0.205	0.918	0.123	0.947
Austria	0.000	0.825	–	–
Belgium	0.945	0.907	0.241	0.057
Canada	0.163	0.980	0.100	0.967
Denmark	0.991	0.971	0.252	0.163
Finland	0.079	0.816	0.145	0.980
France	0.974	0.887	0.030	0.048
Greece	0.994	0.992	0.144	0.242
Ireland	0.863	0.605	0.409	0.807
Italy	0.779	0.963	0.615	0.892
Japan	0.190	0.187	0.757	0.894
Netherlands	0.018	0.005	–	–
New Zealand	0.740	0.989	0.384	0.866
Norway	0.198	0.945	0.149	0.911
Portugal	0.989	0.997	0.036	0.132
Spain	0.904	0.982	0.543	0.723
Sweden	0.174	0.999	0.074	0.958
Switzerland	0.198	0.454	0.020	0.710
United Kingdom	0.239	0.062	0.341	0.673
United States	0.357	0.903	0.495	0.944

Note: Numbers are tail areas of tests.

we present the results for the U.S. dollar-based exchange rates, in panel B for the DM-based exchange rates, and in panel C for the Japanese yen-based exchange rates.

The evidence shows that the null hypothesis that $b_k = 1$ cannot be rejected for most countries when U.S. dollar-based exchange rates are used. In particular, the evidence in panel A of Figs. 1–21 indicates that purchasing power parity cannot be rejected for any $k \in [1, 30]$ for 11 countries: Australia, Denmark, Greece, Italy, the Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, and Switzerland. There is also evidence in support of purchasing power parity for Austria and Germany where the null hypothesis that $b_k = 1$ is rejected only for $k \in [28, 30]$, and for Ireland where $b_k = 1$ is rejected only for $k \in [7, 10]$. For Belgium, Japan, and the United Kingdom there are no permanent changes in their relative prices and thus $\mathrm{LRD}_{s,x}$ is undefined.

When DM-based exchange rates are used, the evidence in panel B of Figs. 1–21 suggests that purchasing power parity cannot be rejected for all values of k only for Australia, Portugal, and Switzerland. There is also evidence consistent with purchasing power parity for France for $k = [1, 16]$, for New Zealand for $k = [1, 26]$, and for Greece and the United States for $k = [1, 27]$. The null hypothesis that $b_k = 1$ is rejected for the rest of the countries, when Germany is used as the home country. Moreover, for Austria and the Netherlands there is no purchasing power issue to be tested since there are no permanent changes in their respective relative prices.

Table 3
Unit root and cointegration tests with Japan as the home country

Country	(A) Unit root tests		(B) Cointegration tests, dependent variable	
	s_t	x_t	s_t	x_t
Australia	0.381	0.988	0.011	0.760
Austria	0.137	0.210	0.696	0.755
Belgium	0.820	0.189	0.724	0.476
Canada	0.045	0.679	–	–
Denmark	0.899	0.557	0.036	0.295
Finland	0.006	0.990	–	–
France	0.860	0.226	0.025	0.451
Germany	0.190	0.187	0.757	0.894
Greece	0.798	0.899	0.237	0.127
Ireland	0.327	0.781	0.019	0.529
Italy	0.201	0.703	0.032	0.778
Netherlands	0.304	0.870	0.849	0.988
New Zealand	0.880	0.989	0.286	0.598
Norway	0.087	0.573	0.024	0.438
Portugal	0.907	0.991	0.087	0.416
Spain	0.311	0.815	0.079	0.752
Sweden	0.093	0.996	0.026	0.744
Switzerland	0.008	0.134	–	–
United Kingdom	0.205	0.925	0.236	0.906
United States	0.287	0.000	–	–

Note: Numbers are tail areas of tests.

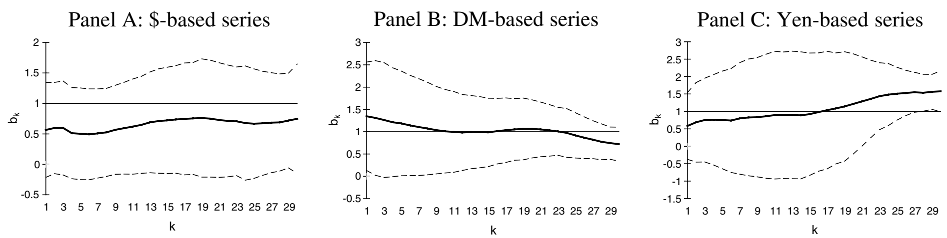

Fig. 1. LRD for Australia.

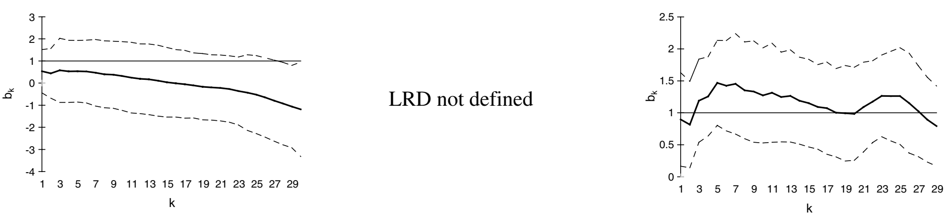

Fig. 2. LRD for Austria.

Finally, when Japan is used as the home country, the results in panel C indicate that purchasing power parity cannot be rejected for any value of k for Australia,

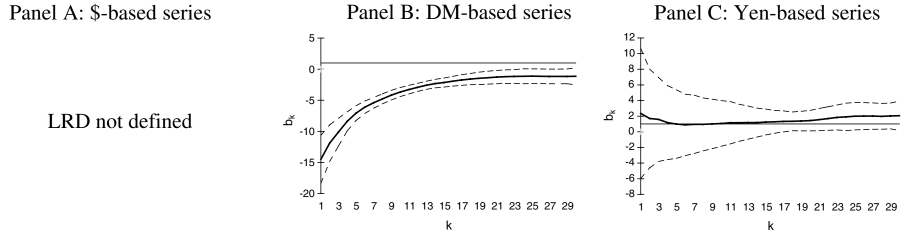

Fig. 3. LRD for Belgium.

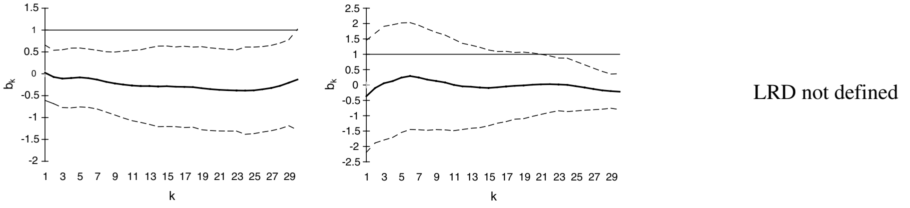

Fig. 4. LRD for Canada.

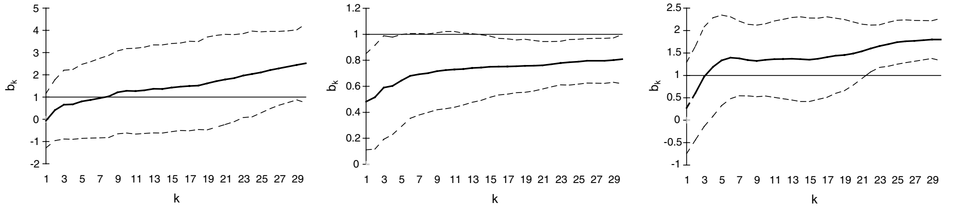

Fig. 5. LRD for Denmark.

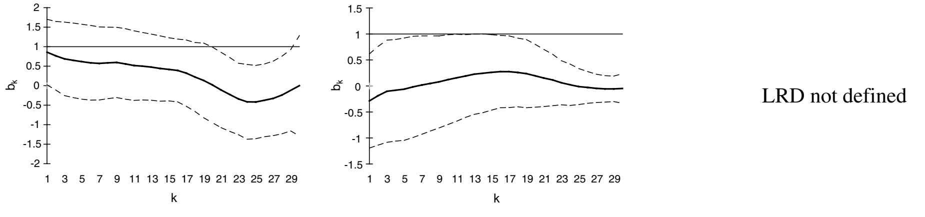

Fig. 6. LRD for Finland.

Austria, Belgium, Finland, Greece, Ireland, Italy, the Netherlands, Norway, Spain, and Sweden. Moreover, for New Zealand and Germany, the $b_k = 1$ null is rejected only for $k = 1$ and $k \in [29, 30]$, respectively. There is also evidence consistent with purchasing power parity for Denmark and France for $k = [1, 21]$ and for the United Kingdom for $k = [1, 18] \cup [27, 30]$.

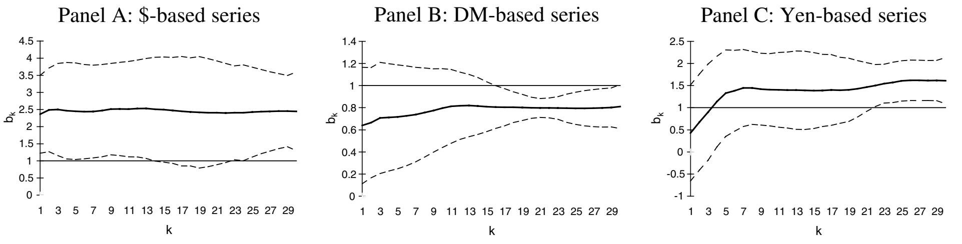

Fig. 7. LRD for France.

Fig. 8. LRD for Germany.

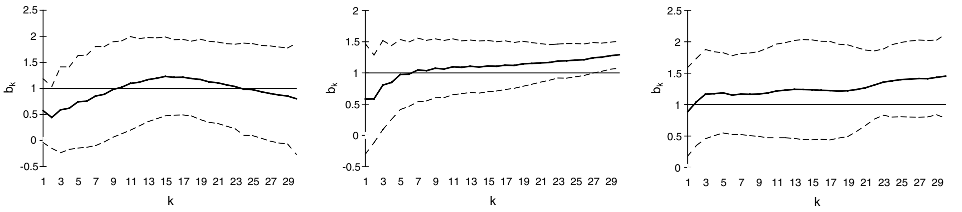

Fig. 9. LRD for Greece.

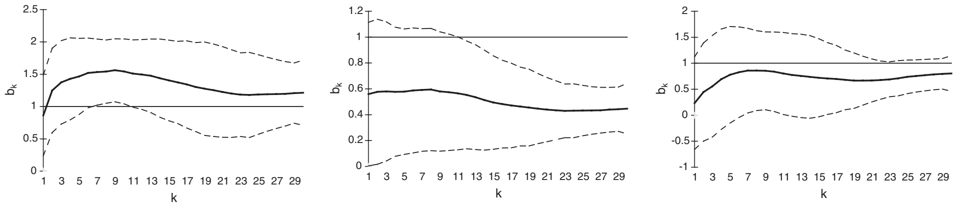

Fig. 10. LRD for Ireland.

5. The power of the LRD tests

The confidence bands in Figs. 1–21 seem extremely wide (especially for large values of k, due to problems with degrees of freedom), suggesting that inference based on these bands has low power. In fact, the long-horizon regression tests of the theory

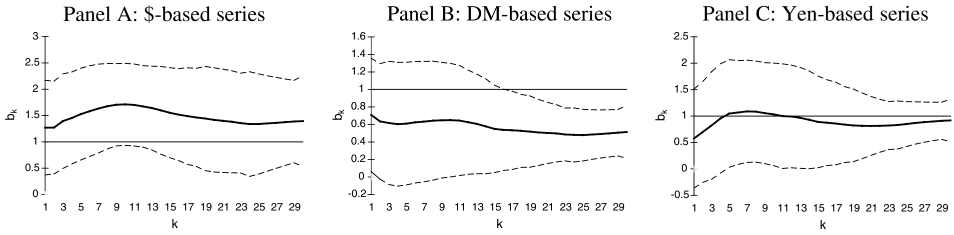

Fig. 11. LRD for Italy.

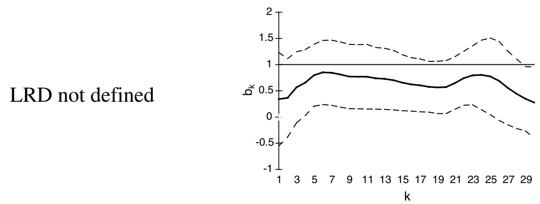

Fig. 12. LRD for Japan.

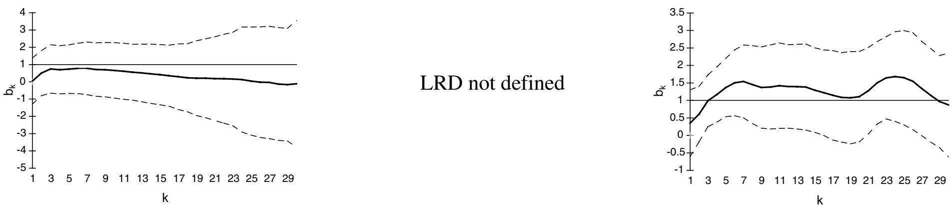

Fig. 13. LRD for the Netherlands.

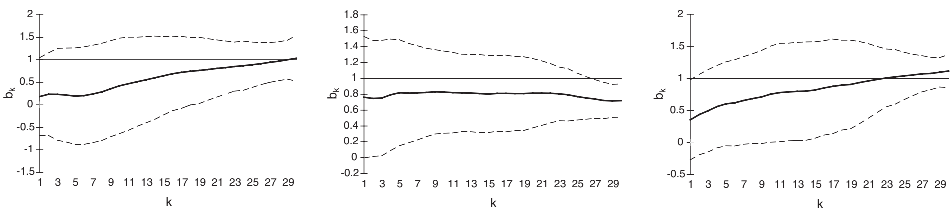

Fig. 14. LRD for New Zealand.

of purchasing power parity that we used in this paper relate to long-horizon regression tests of asset return predictability, and as has been argued by Hodrick (1992), Nelson and Kim (1991), Kirby (1997), Kilian (1999), and Campbell (2001), among others, long-horizon regression tests have substantial size distortions. In fact, Coe and Nason (2004) examine this issue (using Monte Carlo experiments) in the context of long-run monetary neutrality, and show that long-horizon regression tests of

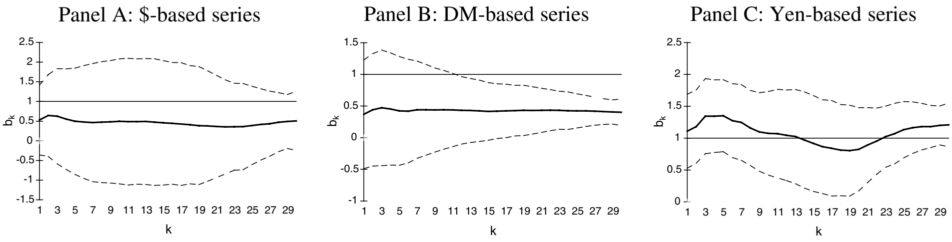

Fig. 15. LRD for Norway.

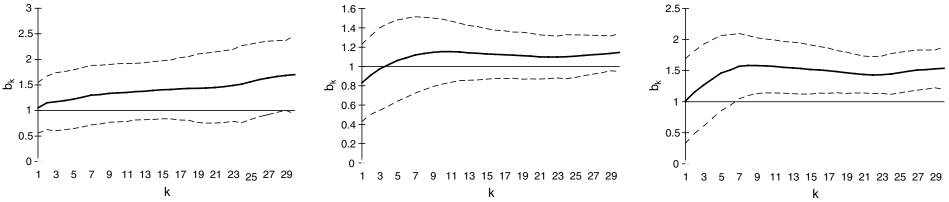

Fig. 16. LRD for Portugal.

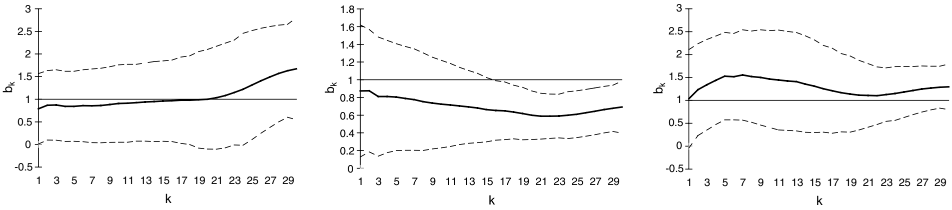

Fig. 17. LRD for Spain.

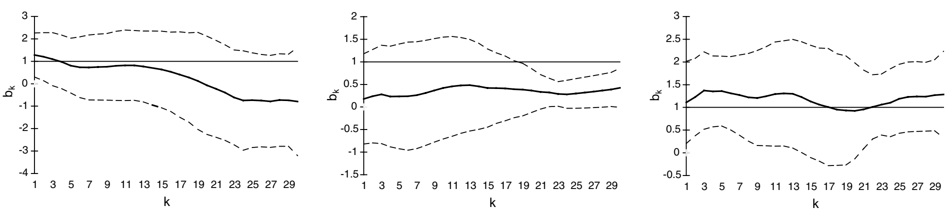

Fig. 18. LRD for Sweden.

long-run monetary neutrality are uninformative because of large-size distortions and that size-adjusted critical values do not offer greater power.

Because we have given PPP the status of the null hypothesis in a test with (apparently) low power, when purchasing power parity is not rejected, only alternatives that have low type II error probability can be ruled out. To separate those alternatives to purchasing power parity that are inconsistent with the data, in this section we

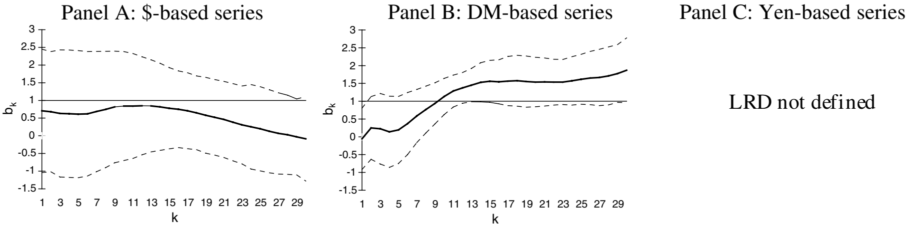

Fig. 19. LRD for Switzerland.

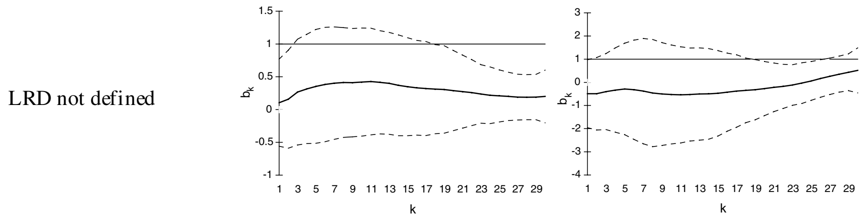

Fig. 20. LRD for the United Kingdom.

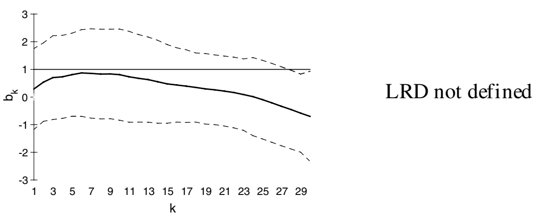

Fig. 21. LRD for the United States.

follow Coe and Nason (2003) and use the inverse power function (IPF) of Andrews (1989) to provide information about deviations from the null hypothesis (of purchasing power parity) when the LRD tests fail to reject the null at a given level of significance.

In Table 4 we report ordinary least squares estimates of b_k, tests of PPP, and Andrews (1989) IPFs, at forecast horizons $k = 10$, 15, 20, 25, and 30, for dollar-based, DM-based, and yen-based exchange rates. An asterisk next to a b_k value indicates rejection of the null hypothesis that $b_k = 1$; the test statistic is the t-ratio of $\hat{b}_k$ with T/k degrees of freedom, where T is the sample size. Clearly, the null hypothesis that $b_k = 1$ is rejected (at the 5% level and for most values of k) with the dollar-based series for Canada and France, the DM-based series for Austria, Belgium, Denmark, Finland, Ireland, Italy, the Netherlands, Norway, Spain, Sweden, and the U.K., and the yen-based series for Canada and Portugal.

The IPF bounds in Table 4 are for low probability of Type II error, $1 \pm b_{k,0.05}$, and for high probability of Type II error, $1 \pm b_{k,0.50}$, where $b_{k,0.05} = \lambda_{1,0.05}(0.95)\hat{\sigma}_{b_k}$ and

Table 4
CPI-Based estimates of b_k, Tests of PPP, and IPF bounds

Country		\$-based series					DM-based series					Yen-based series				
		$k=10$	$k=15$	$k=20$	$k=25$	$k=30$	$k=10$	$k=15$	$k=20$	$k=25$	$k=30$	$k=10$	$k=15$	$k=20$	$k=25$	$k=30$
Australia	$\hat{b}_k$	0.59	0.72	0.75	0.67	0.75	1.01	0.98	1.06	0.92	0.72	0.86	0.92	1.21	1.51	1.58
	$1-b_{k,0.05}$	−0.22	−0.33	−0.35	−0.17	−0.02	−0.52	−0.17	0.08	0.34	0.57	−1.88	−1.68	−1.02	0.01	0.34
	$1+b_{k,0.05}$	2.22	2.33	2.35	2.17	2.02	2.52	2.17	1.92	1.66	1.43	3.88	3.68	3.02	1.99	1.66
	$1-b_{k,0.50}$	0.34	0.28	0.27	0.36	0.45	0.18	0.36	0.50	0.64	0.77	−0.57	−0.46	−0.10	0.46	0.64
	$1+b_{k,0.50}$	1.66	1.72	1.73	1.64	1.55	1.82	1.64	1.50	1.36	1.23	2.57	2.46	2.10	1.54	1.36
Austria	$\hat{b}_k$	0.32	0.04	−0.20	−0.52	−1.19*	−0.09*	−0.02*	−0.04*	−0.16*	−0.24*	1.27	1.15	0.98	1.26	0.70
	$1-b_{k,0.05}$	−1.52	−1.40	−1.07	−1.29	–	–	–	–	–	–	−0.20	−0.04	−0.02	0.02	0.21
	$1+b_{k,0.05}$	3.52	3.40	3.07	3.29	–	–	–	–	–	–	2.20	2.04	2.02	1.98	1.79
	$1-b_{k,0.50}$	−0.37	−0.30	−0.13	−0.24	–	–	–	–	–	–	0.35	0.43	0.45	0.47	0.57
	$1+b_{k,0.50}$	2.37	2.30	2.13	2.24	–	–	–	–	–	–	1.65	1.57	1.55	1.53	1.43
Belgium	$\hat{b}_k$	−0.19	0.75	1.65	2.57	2.74	−3.69*	−2.14*	−1.37*	−1.12*	−1.15*	1.06	1.25	1.44	1.99	2.06
	$1-b_{k,0.05}$	−1.65	−1.33	−1.37	−1.93	−3.24	–	–	–	–	–	−3.73	−1.53	−0.84	−1.28	−1.13
	$1+b_{k,0.05}$	3.65	3.33	3.37	3.93	5.24	–	–	–	–	–	5.73	3.53	2.84	3.28	3.13
	$1-b_{k,0.50}$	−0.44	−0.27	−0.29	−0.60	−1.30	–	–	–	–	–	−1.57	−0.38	0.00	−0.24	−0.16
	$1+b_{k,0.50}$	2.44	2.27	2.29	2.60	3.30	–	–	–	–	–	3.57	2.38	2.00	2.24	2.16
Canada	$\hat{b}_k$	−0.25*	−0.28*	−0.35*	−0.38*	−0.13	0.08	−0.09	0.01	−0.04*	−0.22*	−0.03	−0.50*	−0.59*	0.02	0.29*
	$1-b_{k,0.05}$	–	–	–	–	−0.31	−1.50	−0.88	−0.45	–	–	−1.57	–	–	−0.31	–
	$1+b_{k,0.05}$	–	–	–	–	2.31	3.50	2.88	2.45	–	–	3.57	–	–	2.31	–
	$1-b_{k,0.50}$	–	–	–	–	0.29	−0.36	−0.02	0.21	–	–	−0.40	–	–	0.29	–
	$1+b_{k,0.50}$	–	–	–	–	1.71	2.36	2.02	1.79	–	–	2.40	–	–	1.71	–
Denmark	$\hat{b}_k$	1.28	1.43	1.71	2.11	2.52	0.72	0.75*	0.76*	0.79*	0.8*	1.35	1.35	1.49	1.74*	1.80*
	$1-b_{k,0.05}$	−2.07	−2.02	−1.90	−1.37	−1.01	0.52	–	–	–	–	−0.30	−0.42	0.01	–	–
	$1+b_{k,0.05}$	4.07	4.02	3.90	3.37	3.01	1.48	–	–	–	–	2.30	2.42	1.99	–	–
	$1-b_{k,0.50}$	−0.67	−0.64	−0.58	−0.29	−0.09	0.74	–	–	–		0.29	0.23	0.46	–	–
	$1+b_{k,0.50}$	2.67	2.64	2.58	2.29	2.09	1.26	–	–	–	–	1.71	1.77	1.54	–	–
Finland	$\hat{b}_k$	0.56	0.42	0.02*	−0.42*	0.00	0.12*	0.26*	0.20*	−0.0*	−0.05*	0.71	0.67	0.36	0.37	0.48
	$1-b_{k,0.05}$	−0.46	−0.23	–	–	−0.46	–	–	–	–	–	−2.53	−2.17	−0.95	0.05	0.24
	$1+b_{k,0.05}$	2.46	2.23	–	–	2.46	–	–	–	–	–	4.53	4.17	2.95	1.95	1.76
	$1-b_{k,0.50}$	0.21	0.33	–	–	0.21	–	–	–	–	–	−0.92	−0.72	−0.06	0.49	0.59
	$1+b_{k,0.50}$	1.79	1.67	–	–	1.79	–	–	–	–	–	2.92	2.72	2.06	1.51	1.41

Table 4 (*continued*)

Country		$-based series					DM-based series					Yen-based series				
		$k=10$	$k=15$	$k=20$	$k=25$	$k=30$	$k=10$	$k=15$	$k=20$	$k=25$	$k=30$	$k=10$	$k=15$	$k=20$	$k=25$	$k=30$
France	$\hat{b}_k$	2.52*	2.50	2.41	2.42*	2.45*	0.80	0.81	0.80*	0.79*	0.81	1.41	1.39	1.43	1.61*	1.61*
	$1-b_{k,0.05}$	–	−1.34	−1.21	–	–	0.42	0.66	–	–	0.78	−0.32	−0.31	0.11	–	–
	$1+b_{k,0.05}$	–	3.34	3.21	–	–	1.58	1.34	–	–	1.22	2.32	2.31	1.89	–	–
	$1-b_{k,0.50}$	–	−0.27	−0.20	–	–	0.69	0.81	–	–	0.88	0.28	0.29	0.51	–	–
	$1+b_{k,0.50}$	–	2.27	2.20	–	–	1.31	1.19	–	–	1.12	1.72	1.71	1.49	–	–
Germany	$\hat{b}_k$	0.81	0.48	0.26	−0.09	−0.70*						0.77	0.66	0.57	0.78	0.28*
	$1-b_{k,0.05}$	−1.67	−1.17	−0.77	−0.85	–						0.01	0.18	0.29	0.05	–
	$1+b_{k,0.05}$	3.67	3.17	2.77	2.85	–						1.99	1.82	1.71	1.95	–
	$1-b_{k,0.50}$	−0.45	−0.18	0.04	0.00	–						0.46	0.56	0.61	0.48	–
	$1+b_{k,0.50}$	2.45	2.18	1.96	2.00	–						1.54	1.44	1.39	1.52	–
Greece	$\hat{b}_k$	1.02	1.23	1.12	0.97	0.80	1.06	1.11	1.15	1.21	1.29*	1.19	1.24	1.24	1.40	1.45
	$1-b_{k,0.05}$	−0.44	−0.15	−0.09	−0.15	−0.21	0.26	0.37	0.53	0.65	–	−0.15	−0.21	0.06	0.23	0.24
	$1+b_{k,0.05}$	2.44	2.15	2.09	2.15	2.21	1.74	1.63	1.47	1.35	–	2.15	2.21	1.94	1.77	1.76
	$1-b_{k,0.50}$	0.22	0.37	0.41	0.38	0.34	0.60	0.66	0.75	0.81	–	0.37	0.34	0.49	0.58	0.59
	$1+b_{k,0.50}$	1.78	1.63	1.59	1.62	1.66	1.40	1.34	1.25	1.19	–	1.63	1.66	1.51	1.42	1.41
Ireland	$\hat{b}_k$	1.54*	1.41	1.25	1.19	1.21	0.57	0.49*	0.45*	0.43*	0.45*	0.84	0.72	0.67	0.74	0.80
	$1-b_{k,0.05}$	–	0.05	0.00	0.20	0.43	0.27	–	–	–	–	−0.24	−0.15	0.34	0.58	0.61
	$1+b_{k,0.05}$	–	1.95	2.00	1.80	1.57	1.73	–	–	–	–	2.24	2.15	1.66	1.42	1.39
	$1-b_{k,0.50}$	–	0.48	0.45	0.57	0.69	0.60	–	–	–	–	0.33	0.38	0.64	0.77	0.79
	$1+b_{k,0.50}$	–	1.52	1.55	1.43	1.31	1.40	–	–	–	–	1.67	1.62	1.36	1.23	1.21
Italy	$\hat{b}_k$	1.71	1.56	1.42	1.34	1.39	0.65	0.55	0.51*	0.48*	0.51*	1.03	0.89	0.81	0.85	0.92
	$1-b_{k,0.05}$	−0.26	−0.30	−0.39	−0.24	0.02	−0.05	0.24	–	–	–	−0.57	−0.32	0.13	0.45	0.55
	$1+b_{k,0.05}$	2.26	2.30	2.39	2.24	1.98	2.05	1.76	–	–	–	2.57	2.32	1.87	1.55	1.45
	$1-b_{k,0.50}$	0.31	0.29	0.25	0.33	0.47	0.43	0.59	–	–	–	0.15	0.28	0.53	0.70	0.75
	$1+b_{k,0.50}$	1.69	1.71	1.75	1.67	1.53	1.57	1.41	–	–	–	1.85	1.72	1.47	1.30	1.25
Japan	$\hat{b}_k$	−0.01	−0.38	−0.85*	−0.61	−0.24	0.77	0.66	0.57	0.78	0.28*					
	$1-b_{k,0.05}$	−0.97	−1.12	–	−2.02	−1.64	0.01	0.18	0.29	0.05	–					
	$1+b_{k,0.05}$	2.97	3.12	–	4.02	3.64	1.99	1.82	1.71	1.95	–					
	$1-b_{k,0.50}$	−0.07	−0.15	–	−0.64	−0.44	0.46	0.56	0.61	0.48	–					
	$1+b_{k,0.50}$	2.07	2.15	–	2.64	2.44	1.54	1.44	1.39	1.52	–					

Netherlands	$\hat{b}_k$	0.66	0.41	0.21	0.04	−0.12	0.15*	0.25*	0.31*	0.35*	0.36*	1.38	1.29	1.11	1.65	0.87
	$1-b_{k,0.05}$	−1.60	−1.65	−2.20	−3.07	−3.15	–	–	–	–	–	−0.94	−0.84	−0.81	−0.75	−0.68
	$1+b_{k,0.05}$	3.60	3.65	4.20	5.07	5.15	–	–	–	–	–	2.94	2.84	2.81	2.75	2.68
	$1-b_{k,0.50}$	−0.41	−0.44	−0.74	−1.21	−1.26	–	–	–	–	–	−0.06	0.00	0.01	0.05	0.09
	$1+b_{k,0.50}$	2.41	2.44	2.74	3.21	3.26	–	–	–	–	–	2.06	2.00	1.99	1.95	1.91
New Zealand	$\hat{b}_k$	0.42	0.65	0.79	0.89	1.03	0.83	0.80	0.81	0.77	0.72*	0.76	0.82	0.94	1.04	1.12
	$1-b_{k,0.05}$	−0.72	−0.33	0.04	0.35	0.42	0.16	0.26	0.39	0.62	–	−0.22	−0.15	0.10	0.52	0.71
	$1+b_{k,0.05}$	2.72	2.33	1.96	1.65	1.58	1.84	1.74	1.61	1.38	–	2.22	2.15	1.90	1.48	1.29
	$1-b_{k,0.50}$	0.07	0.28	0.48	0.64	0.69	0.54	0.60	0.67	0.79	–	0.34	0.37	0.51	0.74	0.84
	$1+b_{k,0.50}$	1.93	1.72	1.52	1.36	1.31	1.46	1.40	1.33	1.21	–	1.66	1.63	1.49	1.26	1.16
Norway	$\hat{b}_k$	0.49	0.46	0.38	0.39	0.51	0.44	0.41*	0.43*	0.42*	0.40*	1.08	0.92	0.82	1.13	1.21
	$1-b_{k,0.05}$	−1.57	−1.41	−0.97	−0.29	0.15	−0.01	–	–	–	–	−0.06	−0.15	0.08	0.43	0.61
	$1+b_{k,0.05}$	3.57	3.41	2.97	2.29	1.85	2.01	–	–	–	–	2.06	2.15	1.92	1.57	1.39
	$1-b_{k,0.50}$	−0.40	−0.31	−0.07	0.30	0.54	0.45	–	–	–	–	0.42	0.38	0.50	0.69	0.79
	$1+b_{k,0.50}$	2.40	2.31	2.07	1.70	1.46	1.55	–	–	–	–	1.58	1.62	1.50	1.31	1.21
Portugal	$\hat{b}_k$	1.35	1.41	1.44	1.56	1.70	1.15	1.13	1.11	1.11	1.14	1.58*	1.52*	1.45*	1.47*	1.54*
	$1-b_{k,0.05}$	0.08	0.14	0.04	0.04	0.13	0.45	0.60	0.67	0.71	0.78	–	–	–	–	–
	$1+b_{k,0.05}$	1.92	1.86	1.96	1.96	1.87	1.55	1.40	1.33	1.29	1.22	–	–	–	–	–
	$1-b_{k,0.50}$	0.50	0.53	0.48	0.48	0.53	0.70	0.78	0.82	0.84	0.88	–	–	–	–	–
	$1+b_{k,0.50}$	1.50	1.47	1.52	1.52	1.47	1.30	1.22	1.18	1.16	1.12	–	–	–	–	–
Spain	$\hat{b}_k$	0.91	0.96	1.00	1.31	1.67	0.73	0.66	0.60*	0.61*	0.69*	1.47	1.31	1.11	1.18	1.30
	$1-b_{k,0.05}$	−0.38	−0.36	−0.55	−0.57	−0.28	0.20	0.45	–	–	–	−0.71	−0.54	−0.07	0.28	0.44
	$1+b_{k,0.05}$	2.38	2.36	2.55	2.57	2.28	1.80	1.55	–	–	–	2.71	2.54	2.07	1.72	1.56
	$1-b_{k,0.50}$	0.25	0.26	0.16	0.15	0.31	0.57	0.70	–	–	–	0.07	0.16	0.42	0.61	0.69
	$1+b_{k,0.50}$	1.75	1.74	1.84	1.85	1.69	1.43	1.30	–	–	–	1.93	1.84	1.58	1.39	1.31
Sweden	$\hat{b}_k$	0.80	0.64	−0.07	−0.73	−0.79	0.42	0.42	0.36*	0.30*	0.42*	1.25	1.13	0.92	1.19	1.28
	$1-b_{k,0.05}$	−1.49	−1.60	−2.07	−1.74	−1.73	−0.83	−0.32	–	–	–	−0.76	−0.88	−0.48	0.01	−0.08
	$1+b_{k,0.05}$	3.49	3.60	4.07	3.74	3.73	2.83	2.32	–	–	–	2.76	2.88	2.48	1.99	2.08
	$1-b_{k,0.50}$	−0.36	−0.42	−0.67	−0.49	−0.49	0.01	0.28	–	–	–	0.04	−0.02	0.19	0.46	0.41
	$1+b_{k,0.50}$	2.36	2.42	2.67	2.49	2.49	1.99	1.72	–	–	–	1.96	2.02	1.81	1.54	1.59

Table 4 (*continued*)

Country		$-based series					DM-based series					Yen-based series				
		$k = 10$	$k = 15$	$k = 20$	$k = 25$	$k = 30$	$k = 10$	$k = 15$	$k = 20$	$k = 25$	$k = 30$	$k = 10$	$k = 15$	$k = 20$	$k = 25$	$k = 30$
Switzerland	$\hat{b}_k$	0.84	0.78	0.52	0.20	−0.09	1.13	1.56	1.54	1.62	1.87	1.10	1.07	1.12	1.16	0.92
	$1 - b_{k,0.05}$	−1.49	−0.75	−0.49	−0.55	−0.35	0.18	0.10	0.03	0.09	−0.03	−0.08	0.49	0.40	0.21	0.41
	$1 + b_{k,0.05}$	3.49	2.75	2.49	2.55	2.35	1.82	1.90	1.97	1.91	2.03	2.08	1.51	1.60	1.79	1.59
	$1 - b_{k,0.50}$	−0.36	0.05	0.19	0.16	0.26	0.55	0.51	0.47	0.50	0.44	0.41	0.72	0.67	0.57	0.68
	$1 + b_{k,0.50}$	2.36	1.95	1.81	1.84	1.74	1.45	1.49	1.53	1.50	1.56	1.59	1.28	1.33	1.43	1.32
U.K.	$\hat{b}_k$	0.88	0.49	0.11*	0.10*	0.64	0.42	0.35	0.29*	0.2*	0.20*	−0.52	−0.48	−0.27*	0.06*	0.52
	$1 - b_{k,0.05}$	−0.03	−0.10	–	–	0.27	−0.33	−0.14	–	–	–	−2.46	−1.81	–	–	−0.10
	$1 + b_{k,0.05}$	2.03	2.10	–	–	1.73	2.33	2.14	–	–	–	4.46	3.81	–	–	2.10
	$1 - b_{k,0.50}$	0.44	0.40	–	–	0.61	0.28	0.38	–	–	–	−0.88	−0.53	–	–	0.40
	$1 + b_{k,0.50}$	1.56	1.60	–	–	1.39	1.73	1.62	–	–	–	2.88	2.53	–	–	1.60
U.S.	$\hat{b}_k$						0.81	0.48	0.26	−0.09	−0.70*	−0.01	−0.38	−0.85*	−0.61	−0.24
	$1 - b_{k,0.05}$						−1.67	−1.17	−0.77	−0.85	–	−0.97	−1.12	–	−2.02	−1.64
	$1 + b_{k,0.05}$						3.67	3.17	2.77	2.85	–	2.97	3.12	–	4.02	3.64
	$1 - b_{k,0.50}$						−0.45	−0.18	0.04	0.00	–	−0.07	−0.15	–	−0.64	−0.44
	$1 + b_{k,0.50}$						2.45	2.18	1.96	2.00	–	2.07	2.15	–	2.64	2.44

Note: An asterisk indicates rejection (of the null that $b_k = 1$) at the 5% asymptotic level.

$b_{k,0.50} = \lambda_{1,0.05}(0.50)\hat{\sigma}_{b_k}$, with $\hat{\sigma}_{b_k}$ being the standard error of $\hat{b}_k$. The subscripts of $\lambda_{q,\alpha}(1-\gamma)$ denote the number of restrictions and the significance of the test, and γ is the probability of Type II error. From Andrews (1989, Table 1) we get that $\lambda_{1,0.05}(0.95) = 3.605$ and $\lambda_{1,0.05}(0.50) = 1.96$. In particular, $1 - b_{k,0.05}$ and $1 + b_{k,0.05}$ define the region $\Omega = (-\infty, 1 - b_{k,0.05}] \cup [1 + b_{k,0.05}, +\infty)$ where the probability of Type II error is small (0.05 or less). In this case, when we fail to reject the null, we can say with significance 0.05 that there is evidence against any parameter value in Ω, or in other words that the true value of b_k is $1 - b_{k,0.05} < b_k < 1 + b_{k,0.05}$; the evidence that we have against any parameter value in Ω when we fail to reject, is similar to the evidence against null parameter values when the test does reject.

Similarly, the estimates of $1 - b_{k,0.50}$ and $1 + b_{k,0.50}$ define the region of high probability of Type II error, 0.50 or higher suggested as an obvious focal point by Andrews (1989), $\Psi = [1 - b_{k,0.50}, 1 + b_{k,0.50}]$. In this region, the probability of Type II error is 0.50 or higher, suggesting that the power of the test, defined as $P = 1 - \gamma$, is 0.50 or less; the test provides no evidence for parameter values in region Ψ as one has a better chance of rejecting a false null by tossing a fair coin than by using the test. Clearly, the narrower regions Ω and Ψ are around the tested parameter value, the higher is the power of the test. To illustrate the usefulness of these measures of power, consider the results for Portugal with the DM-based series for $k = 30$. In this case, $\Omega = (-\infty, 0.78] \cup [1.22, +\infty)$ and $\Psi = [0.88, 1.12]$. This means that the IPFs show that $0.78 < b_{30} < 1.22$ with significance level 0.05, but due to the low power there is no evidence that it is $0.88 < b_{30} < 1.12$. If we accept that $b_k \in [0.78, 1.22]$ is not a significant (from a theoretical point of view) deviation from PPP, then we conclude that PPP holds.

The IPFs in Table 4 reveal the low power of some of the long-horizon regression tests of purchasing power parity. Consider, for example, the dollar-based series for Belgium. The estimate of b_k for $k = 30$ is 2.74 and the null hypothesis that $b_k = 1$ cannot be rejected. However, the low and high probability of type II error regions are $\Omega = (-\infty, -3.24] \cup [5.24, +\infty)$ and $\Psi = [-1.30, 3.30]$, meaning that the LRD test cannot rule out not only $b_k = 1$ as the true parameter value (in which case PPP holds), but also any other parameter value in Ψ, including 0 (in which case PPP does not hold). Hence, we can mark as "inconclusive" all the tests in Table 1 where both 0 and 1 belong to Ψ. In those cases that the null cannot be rejected and the tests are not inconclusive, the tighter the bands of regions Ω and Ψ are around 1, the higher the power of the test is (and the more confident we are that PPP holds). For example, the results for Portugal show that with the DM-based series there is very strong evidence in favor of purchasing power parity, for all values of k; Ψ is at its widest, $[0.70, 1.30]$, when $k = 10$, and at its tightest, $[0.88, 1.12]$ when $k = 30$.

Assuming that the IPFs in the area of low power are in the region $[1 \pm 0.4]$ or $\Psi = [0.60, 1.40]$, the LRD test is powerful enough with the dollar-based series for Ireland (for $k = 30$), New Zealand (for $k = 25$, 30), and the United Kingdom (for $k = 15$ and 30). With the DM-based series, the tests that support PPP are powerful for Australia (for $k = 25$, 30), France (for $k = 10$, 15, 30), Greece (for all k except $k = 30$), New Zealand (for $k = 15$, 20, 25), Portugal (for every k), Denmark and

Ireland (for $k = 10$), Japan (for $k = 20$), and Spain (for $k = 15$). Finally, with the yen-based series, the LRD tests that support PPP are powerful in the case of Australia (for $k = 30$), Germany (for $k = 20$), Ireland (for $k = 20$, 25, 30), Italy, New Zealand, Norway, and Spain (for $k = 25$, 30), and Switzerland (for $k = 15$, 20, 30).

Of course, for cut-off values lower than 0.4, we reach different conclusions regarding the power of the long-horizon regression tests. Assuming for example, that the IPFs in the area of low power are in the region $[1 \pm 0.1]$ or $\Psi = [0.90, 1.10]$, then we conclude that the long-horizon regression tests that support PPP have low power for all the bilateral relations considered.

6. Robustness

Although consumer price indices used so far in this paper (due to the lack of alternative output price indices for all 21 countries examined) are less sensitive to exchange rate movements, as Burstein et al. (2002) argue exchange rate fluctuations have a bearing on output prices through the role of imported raw materials and intermediate goods in production. In this section, we investigate the robustness of our results to alternative output price indices, using GDP deflators as well as producer price indices (PPIs) to calculate relative prices.

In particular, we use nominal and real GDP data from the IMF International Financial Statistics to calculate the GDP deflators for those countries for which GDP data exist and then test for purchasing power parity in the same fashion as in Table 4 – the results of these tests are available upon request. Greece, Ireland, and New Zealand are not included in these tests because of data availability problems. Moreover, for a number of countries the deflator-based results are not directly comparable to those in Table 4, because the GDP data are not available since 1973Q1. In particular, GDP data for Belgium begin in 1985Q1, Denmark in 1988Q1, the Netherlands in 1977Q1, New Zealand in 1988Q3, Portugal in 1977Q1, and Sweden in 1980Q1. Based on the evidence using deflator-based relative prices, we reject purchasing power parity with the dollar- and DM-based series for all countries. Moreover, we reject PPP for the yen-based series except for Finland (for $k = 30$), France and Italy (for $k = 25$, 30), Norway and Portugal (for $k = 30$), Spain (for $k = 25$, 30), and Switzerland (for $k = 15$, 30).

Finally, we also perform long-horizon regression tests of purchasing power parity using producer price indices – these results are also available upon request. Finland and Portugal are not included in these tests because of data availability problems, and the data for Belgium begin in 1980Q1, Italy in 1981Q1, and Norway in 1977Q1. With the PPI-based estimates of b_k and the dollar as the base currency, the LRD test provides evidence in support of purchasing power parity with adequate power for Greece (for $k = 25$, 30), Ireland (for $k = 30$) and New Zealand (for $k = 20$, 25, 30). With the DM-based series, the IPFs show that purchasing power parity is supported with enough power for Australia (for $k = 30$), Denmark and Switzerland (for all k except $k = 10$), Belgium, France and Greece (for all k), Ireland (for $k = 10$), New Zealand (for $k = 20$, 25, 30), Spain (for all k except $k = 25$), and Sweden (for

$k = 25, 30$). Finally, with the Japanese yen as the base currency, the LRD tests support purchasing power parity with sufficient power for Australia and Spain (for $k = 25, 30$), Ireland (for all k except $k = 20$), New Zealand (for $k = 20, 25, 30$), Sweden (for $k = 30$) and Switzerland (for all k except $k = 10$).

7. Conclusion

We have tested the purchasing power parity hypothesis using quarterly data (from the IMF International Financial Statistics) for the recent floating exchange rate period (1973:1–1998:4) for 21 OECD countries, using U.S. dollar-based, DM-based, and Japanese yen-based exchange rates. In doing so, we have used the long-run regression approach of Fisher and Seater (1993) and also investigated the power of the long-run regression tests using Andrews' (1989) inverse power functions.

We have demonstrated that the asymptotic power of the long-horizon regression tests is low, and found weak evidence consistent with purchasing power parity, although the evidence is stronger when producer prices are used. These results are consistent with those empirical studies (mentioned in Section 1) covering the major industrial countries over the recent floating exchange rate period (using tests other than the ones we have conducted), which generally fail to find support for purchasing power parity. They are, however, in contrast to those studies (mentioned in the introduction) covering different groups of countries, as well as those covering periods of long duration, or country pairs experiencing large differentials in price movements.

In using the Fisher and Seater (1993) long-run regression approach, we assumed that exchange rates have no long-run effects on relative prices. Although consumer price indices are less sensitive to exchange rate movements, as already noted exchange rate fluctuations have a bearing on output prices through the role of imported raw materials and intermediate goods in production. Hence, investigating the robustness of our results to alternative testing methodologies such as, for example, the King and Watson (1997) approach which allows both variables to be determined endogenously, is an area for potentially productive future research.

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