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The re-emergence of PPP in the 1990s

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Abstract

In this paper we investigate purchasing power parity (PPP) in a panel with 17 countries for the period 1972 through 1996. The novel feature of our panel methodology is that results are invariant to the choice of a benchmark on numeraire currency. In the panel we allow individual country effects in the relation between prices and exchange rates. In this way we can identify the currency pairs for which PPP holds or does not hold. We conclude that there is substantive evidence for PPP, although not to the same extent for every currency. Evidence in favor of PPP is strongest for many exchange rates relative to the Dmark, and weakest for the Japanese yen. For this currency a trend-like variable, like productivity growth, is missing.

JEL classifications: F31; G15

The theory of purchasing power parity (PPP) is one of the central tenets in international economics. When exchange rates started to float worldwide in 1973, it was widely believed that PPP would provide an accurate description of movements in exchange rates. The years of very high exchange rate volatility quickly destroyed that idea. The demise of PPP was also confirmed through formal econometric tests in the late 1970s and early 1980s. At the beginning of the 1980s it seemed as if the theory of PPP had collapsed completely.

In the past decade, tests of PPP have often taken the form of unit root tests of the real exchange rates. As is well-known (see for example the review by Edison et

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al., 1997) these tests suffer from extreme low power and this might have been the reason that the unit root in real exchange rates was hardly ever rejected. Researchers reacted to the low power problem by using long-horizon data like 200 or more years of, for instance, the UK pound/US dollar exchange rate. Well-known studies with long data spans are Edison (1987), and Lothian and Taylor (1996). These studies typically find that real exchange rates exhibit slow but significant mean-reversion suggesting that PPP might be valid in the long-run.

Another way to circumvent the low power of the traditional tests has been the use of panel data. Well-known examples in this vein are Abauf and Jorion (1990), and recently Jorion and Sweeney (1996) and Frankel and Rose (1996). Pooling data for different exchange rates against the US dollar, these studies generally find relatively stronger evidence in favor of PPP. On the basis of long-horizon and panel data results, it is generally perceived that PPP has risen from its ashes in the past few years. In his review of the empirical literature Rogoff (1996, pp. 657–658) concludes:

> Overall, while there are some limitations to both the long-horizon and cross-section results on convergence to PPP, the recent literature has reached a surprising degree of consensus: PPP deviations tend to damp out, but only at the slow rate of roughly 15% per annum.

Recently, however, the panel results of PPP have been questioned in a study by O'Connell (1997), who showed that the standard practice of calculating all real exchange rates relative to the US dollar leads to cross-sectional dependence in panel data. Adjusting for this problem makes it much more difficult again to reject the random walk in real exchange rates.

In this paper we consider PPP using a panel data methodology that explicitly deals with the numeraire effect that causes the cross-sectional dependence. Our panel model extends the four country model of Koedijk and Schotman (1990). The parameter estimates in the PPP equations are invariant to choice of numeraire currency.

The focus of the paper is on another assumption that is typical in the panel literature. The power of the panel studies comes from the assumption that PPP holds equally well for every currency. This assumption leads to equality of all mean reversion parameters in the Abauf and Jorion (1990) and Frankel and Rose (1996) studies. Individual currency effects are treated as constant terms (fixed or random effects), while the slope coefficients are equal across equations. An intermediate position is that (long-run) PPP holds for some currency pairs, but not for others. For each country in our panel we investigate whether the value of its currency moves proportionally to the price level (or inflation) in that country.

The difference between our approach and other panel data studies can be explained by the following example. Consider a panel with three countries: the United States, Germany and the Netherlands. Suppose that PPP holds between the mark and the guilder, but not between the dollar and these currencies. No matter which currency is used as the numeraire, PPP is always rejected in this panel. With the dollar as the numeraire the rejection is correct, but with the guilder as the numeraire the rejection is due only to the dollar/guilder rate, and not the mark/guilder rate. Nevertheless, in small samples it would appear that rejection of PPP is stronger with the US dollar as the numeraire than with the guilder as the numeraire. This is exactly the conclusion of Papell and Theodoridis (1998) who compare panel unit root tests with the mark and the dollar as alternative numeraires. In our approach we estimate the price parameters simultaneously for all currencies. In the three-country example we would find that prices and currency values move proportionally in the Netherlands and Germany, while they do not in the United States.

The plan of this paper is as follows. In Sec. 1, we discuss our numeraire invariant panel methodology, while Sec. 2 contains the results. Section 3 concludes.

1. Methodology

The general framework for empirical tests of absolute PPP is to compare consumer price indices expressed in a common numeraire currency. This absolute consumption-based PPP relates the logarithm of the exchange rate between currencies i and j to the logarithm of the consumer price indices in countries i and j:

$$\begin{aligned} q_{ij}(t) &= s_{ij}(t) + p_i(t) - p_j(t) \\ &= (c_i - c_j) + \beta_i p_i(t) - \beta_j p_j(t) + \nu_{ij}(t), \end{aligned} \tag{1}$$

where s_{ij} is the logarithm of the nominal exchange rate (the price of currency j in units of currency i) q_{ij} is the real exchange rate, and p_i denotes the logarithm of the consumer price index for country i. Absolute PPP is said to hold in the long run if the following three conditions are fulfilled: (i) $\beta_i = \beta$ (symmetry), (ii) $\beta = 0$ (proportionally) and (iii) the error term $\nu_{ij}(t)$ is stationary. Most empirical tests simply impose the first two conditions. Under these restrictions the null hypothesis most often tested is that the real exchange rate $q_{ij}(t)$ contains a unit root, against the alternative hypothesis that the real exchange rate is stationary.[1]

The emphasis in this paper is, in contrast, on these first two conditions. Equation (1) is estimated simultaneously as a system of N equations for N exchange rates ($i = 1,\ldots,N$) against the common numeraire currency $j = 0$. We test the null hypothesis $\beta_i = 0$ for each of the $N + 1$ currencies under the maintained hypothesis that the error term $\nu_{ij}(t)$ is stationary (although probably highly autocorrelated). To minimize the risk of spurious regression we consider an augmented version of (Eq. (1)) with a linear trend included:

[1] Incidentally, this null hypothesis is contrary to the usual methodology of hypothesis testing. The theory should hold under the null, and be rejected under the alternative. In the unit root tests the hypotheses are reversed. See Schotman and van Dijk (1991) for a Bayesian analysis of the unit root hypothesis in real exchange rates, in which null and alternative are treated symmetrically.

$$q_{ij}(t) = (c_i - c_j) + (\delta_i - \delta_j)t + \beta_i p_i(t) - \beta_j p_j(t) + \nu_{ij}(t). \quad (2)$$

The constant terms c_j and trend coefficients δ_j are not identified. Without loss of generality we can normalize these parameters using $c_0 = 0$ and $\delta_0 = 0$. Under the null hypothesis of long-run PPP the remaining trend coefficients $\delta_i (i = 1,\ldots,N)$ should be equal to zero, as any other trending explanatory variable.

For ease of interpretation we report the results for the transformed, but equivalent representation

$$q_{i0}(t) = (\tilde{c}_i - \tilde{c}_0) + \left(\tilde{\delta}_i - \tilde{\delta}_0\right)t + \beta_i \tilde{p}_i(t) - \beta_0 \tilde{p}_0(t) + u_{i0}(t), \quad (3)$$

where $\tilde{p}_j(t) = p_j(t) - \psi_{0j} - \psi_{1j}t$ is the detrended price series defined as the residuals of the regression of $p_j(t)$ on a constant and a linear trend, so that

$$\tilde{c}_i - \tilde{c}_0 = c_i - c_0 + \beta_i \psi_{0i} - \beta_0 \psi_{00},$$
$$\tilde{\delta}_i - \tilde{\delta}_0 = \delta_i - \delta_0 + \beta_i \psi_{1i} - \beta_0 \psi_{10}.$$

Detrending prices on the right hand side of (Eq. (3)) will not affect the estimates of β_i, but has the advantage that prices have been orthogonalized from the trend component. In the extended model, any effect of $\tilde{p}_i(t)$ on the real exchange rate is due to price variability and does not come about through a missing trend in the real exchange rate. In the transformed model, PPP — the null hypothesis — remains unchanged.

The levels tests have some serious defects. Because exchange rates and prices might have unit roots, the test statistics may not have a standard asymptotic distribution. Even if the explanatory variables do not have an exact unit root, the results in Stock (1996) indicate that tests based on standard asymptotic theory could be unreliable. To circumvent unit root and spurious regression problems, we also consider the hypothesis of relative PPP. The relative version of PPP requires that the percentage change in the bilateral exchange rate equals the inflation differential between the domestic and foreign country. The hypothesis does not specify the horizon over which inflation differentials and exchange rate changes should be equal. Taking k-period differences of (Eq. (1)), the test equation for relative PPP becomes:

$$\Delta_k q_{ij}(t) = \beta_i \Delta_k p_i(t) - \beta_j \Delta_k p_j(t) + \nu_{ij}(t), \quad (4)$$

where Δ_k denotes the k-period difference operator $\Delta y(t) = y(t) - y(t-k)$, and where the error term $\nu_{ij}(t)$ is possibly autocorrelated due to overlapping observations. For model (Eq. (2)) with trends the differencing operation leads to an augmented version of (Eq. (4)) with a constant term.

Estimation and testing is carried out on a full panel of $N + 1$ currencies. Suppose we have a sample with currencies numbered $i = 0,1\ldots,N$. For the actual estimation we only need data relative to one particular numeraire currency, say currency 0. Simply subtracting the equations for exchange rates q_{i0} and q_{j0} yields the implied regression model for the cross exchange rate q_{ij}, which automatically has explanatory variables $\Delta_k p_i$ and $\Delta_k p_j$, while $\Delta_k p_0$ drops out.

Since all data in the regression are expressed in the same numeraire currency 0, the error terms $\nu_{i0}(t)$ $(i = 1,\ldots,N)$ are likely to be positively correlated due to the strong common numeraire effect. This implies that, although the cross-sectional ordinary least squares (OLS) estimator is consistent, it will not be efficient. More efficient estimates can be obtained by applying generalized least squares (GLS), which requires assumptions about the error terms $\nu_{i0}(t)$. As a model for the cross-sectional dependence we assume the decomposition of Mahieu and Schotman (1994):

$$\nu_{i0}(t) = \nu_i(t) - \nu_0(t). \tag{5}$$

The decomposition states that the error term in the exchange rate equation is the difference between an error term $\nu_i(t)$ for country i and an error term for the numeraire country 0, which appears in N equations of the system. In the panel literature this specification is referred to as the random time effects model (see Baltagi, 1995). We assume that the country-specific shocks are mutually uncorrelated and have a common variance $\sigma^2/2$. Under these assumptions the covariance matrix for the vector $\nu_0 = (\nu_{10},\ldots \nu_{N0})'$ takes the form:

$$\Sigma = \tfrac{1}{2}\sigma^2(\mathbf{I} + \iota\iota'), \tag{6}$$

with $\mathbf{I}$ the $(N \times N)$ identity matrix and ι the $(N \times 1)$ vector of ones. This covariance structure imposes the conditions that all exchange rates have equal variance and that the correlation between exchange rate changes is $1/2$. Since Σ is completely specified, the GLS estimator is directly applicable.

The covariance structure in (Eq. (6)) not only deals with the positive cross-sectional correlations, but as Koedijk and Schotman (1990) show, it also ensures that all results are completely invariant with respect to the choice of the numeraire currency. Whether we express all exchange rates against the US dollar, the Japanese yen (JPY), the Deutschmark (DEM) or any other currency, the estimates of the parameters β_j and δ_j will be identical.

To compute standard errors we use the Newey–West procedure. Application of the Newey–West estimator provides standard errors that are robust against heteroskedasticity and autocorrelation in the errors. Also, since the Newey–West estimator does not employ the cross sectional covariance structure (Eq. (6)), it is robust against possible misspecification of the covariance matrix Σ in the panel model. Even if the cross equation covariance structure is more complicated than in (Eq. (6)), the GLS estimator of β remains consistent, although no longer efficient.

2. Results

The data are quarterly data and cover a period of 24 years (1973:1–1996:3). Nominal exchange rates, consumer prices and real expenditures have been collected for 17 countries: Australia, Austria, Canada, Finland, France, Germany, Italy, Japan, the Netherlands, Norway, Spain, Sweden, Switzerland, United Kingdom, Belgium, Denmark and the United States. The nominal exchange rate and the consumer price index (CPI) are available from the *International Financial Statistics* (*IFS*) tape (lines ae and 64).

All series are converted to logarithms. Real exchange rates against the dollar are constructed as $s_{i0} - p_0 + p_i$, where s_{i0} is the logarithm of the nominal exchange rate against the US dollar, p_i denotes the logarithm of the consumer price index in country i, and p_0 denotes the logarithm of consumer price index for the United States.

Table 1 refers to the system of levels regressions (Eq. (1)). Under the null hypothesis of PPP all the β_i must be equal to zero. The pooled estimate shows that for these 17 countries taken together we cannot reject the PPP hypothesis, since the level of the real exchange rate is not related to relative prices. However, the pooling restriction $\beta_i = \beta$, for all $i = 0, \ldots N$, is rejected. A Wald test based on the unrestricted parameter estimates rejects with a P-value less than 0.001. The parameter estimates for the individual currencies show much variation, with especially Japan and Switzerland being very far from zero, and significantly so. The PPP hypothesis is rejected for five countries: Belgium, Canada, Japan, Sweden, and Switzerland.

Table 2 repeats the PPP test with a linear trend added to the explanatory variables as in (Eq. (2)). With a trend included, the hypothesis $\beta_i = 0$ is rejected much more often. For example, for the United States the parameter estimate $\beta_{US} = -0.99$ implies that the coefficient on the nominal US dollar exchange rates would be only 0.01. In other words, the covariance between the nominal dollar exchange rate and the US price level in deviation of a trend is almost zero, so that nominal US dollar exchange rates are not affected at all by consumer prices. The same holds for Sweden, and to a lesser extent for Canada, Finland, and Norway. The countries for which we do not reject the hypothesis $\beta_i = 0$ are Belgium, Denmark, France, Germany and Italy. The parameter β_i is negative for all currencies, indicating that exchange rates do not fully adjust to (permanent) price increases.

When prices do not explain the trends in the nominal exchange rates, the linear trend terms must account for them. To interpret the trend coefficient $\tilde{\delta}_i$ consider for example Australia. The parameter estimate of 0.24 relative to the US dollar means that the real Australian dollar has been depreciating against the US dollar at a rate of approximately 1% per year (0.24 × 4 quarters), but at the much slower rate of $4 \times (0.24 - 0.18) \approx 1/4\%$ against the Canadian dollar. At a rate of almost 4% per year (against the US dollar) the Japanese yen is the strongest appreciating currency in the system.

Table 3 reports results for the tests on changes in exchange rates. The results are reported both with and without a constant term in each equation. As for the level regressions the slope coefficients are much closer to zero for most currencies when the constant terms are left out. However, the constants are often jointly significant. This means that some real exchange rate changes have a non-zero drift. For panel

Table 1
Absolute PPP regression

Country	β_j	Country	β_j	Country	β_j
Australia	0.20	France	0.07	Japan	−0.53
	(0.28)		(0.05)		(0.10)
Austria	−0.11	Germany	0.15	Netherlands	0.13
	(0.08)		(0.11)		(0.09)
Belgium	0.15	Italy	−0.01	Norway	0.04
	(0.07)		(0.04)		(0.05)
Canada	0.19	Spain	−0.06	United Kingdom	−0.03
	(0.06)		(0.04)		(0.05)
Denmark	0.01	Sweden	0.15	United States	0.08
	(0.05)		(0.05)		(0.06)
Finland	−0.02	Switzerland	−0.28		
	(0.06)		(0.10)		
Pooled	0.05				
	(0.03)				

Sample period is 1972.I–1996.III, quarterly data. The table reports numeraire invariant regression coefficients β_i $(i = 0,\ldots,N)$ of the effect of the price level on the real exchange rate in the model $q_{ij} = c_i - c_j + \beta_i p_i - \beta_j p_j + \nu_{ij}$, under the identifying restriction $c_0 = 0$. Price index is the CPI. Standard errors have been computed from the spectral density at zero, and are reported in parentheses.

unit root tests of long-run PPP this implies that a trend $\delta_i t$ cannot be omitted from the test regression.

For one quarter differences relative PPP is strongly rejected. For most currencies the slope coefficient is significantly different from zero. For the 12 quarter differences there are only two statistical rejections: Canada and Norway. However, since the standard errors are large, the test might not be very powerful. Equality of all slopes is always rejected, so the pooled estimate should be interpreted with care.

The most remarkable results are for Japan. For the yen the inclusion of a constant term brings the inflation effect much more in line with relative PPP. Since the constant term represents a trend in the real exchange rate, a trend-like variable is clearly missing. But apart from the trend the marginal effect of inflation on the exchange rate is fully consistent with PPP. The trend in the Japanese yen is often attributed to a sharp postwar increase in productivity in Japan. Rogoff (1996) calls the yen/dollar exchange rate the ‘canonical time series example of the Balassa Samuelson effect’ (p. 661). Chinn and Johnston (1996) conduct an extensive panel cointegration study for 14 countries, and conclude that the productivity trend greatly improves the fit for the yen/dollar exchange rate, but much less so for other exchange rates against the US dollar.

Interestingly, we find that many of the slope parameters jump upward at the four-quarter horizon. Notice for instance the shift in the coefficient for the pooled estimate, which increases from −0.43 to −0.10 when we move from a one-quarter to a 1-year horizon. This could mean that some of the statistical difficulties in finding evidence of PPP are due to seasonal measurement errors in price indices.

Table 2
Augmented absolute PPP regression

Country	$\tilde{\beta}_i$	$\tilde{\delta}_i$	Country	$\tilde{\beta}_i$	$\tilde{\delta}_i$
Australia	−0.36	0.24	Japan	−0.26	−0.73
	(0.07)	(0.04)		(0.10)	(0.04)
Austria	−0.34	−0.27	Netherlands	−0.34	−0.02
	(0.16)	(0.05)		(0.11)	(0.05)
Belgium	−0.01	0.01	Norway	−0.49	−0.07
	(0.17)	(0.06)		(0.06)	(0.04)
Canada	−0.69	0.18	Spain	−0.17	−0.33
	(0.22)	(0.02)		(0.08)	(0.05)
Denmark	−0.06	−0.17	Sweden	−0.16	0.12
	(0.08)	(0.05)		(0.17)	(0.04)
Finland	−0.56	−0.15	Switzerland	−0.99	−0.39
	(0.12)	(0.04)		(0.30)	(0.07)
France	−0.06	−0.05	United Kingdom	−0.50	−0.18
	(0.05)	(0.05)		(0.11)	(0.03)
Germany	−0.02	−0.05	United States	−0.99	—
	(0.21)	(0.05)		(0.23)	(—)
Italy	−0.20	−0.18			
	(0.13)	(0.04)			

Sample period is 1972.I–1996.III, quarterly data. The table reports numeraire regression coefficients of the effect of the price level on the real exchange rate in model (Eq. (3)). Price index is the CPI. The δ_t parameters refer to the coefficient on the quarterly trend t under the identifying restriction $\delta_{US} = 0$. Standard errors have been computed from the spectral density at zero.

The price data we use are all seasonally adjusted, but in the regressions we always find a large difference between the one-quarter and 1-year regressions. The estimates for eight and 12 quarters are almost identical, and close to those for four quarters.

Whatever way we run the regressions (levels or differences), we always find that for the group of currencies related to the mark, PPP seems a reasonable hypothesis: there are no trends and the slopes are consistent with PPP. For the US (and Canada) as a pair there does not seem any relation at all between prices and exchange rates. For Japan and Switzerland an additional trend is clearly missing.

3. Conclusion

In the past 20 years the PPP pendulum has swung from total collapse in the early 1980s to complete resurrection in the 1990s. The evidence supporting PPP comes especially from long-horizon and panel data results. Recently, however, the panel results of PPP have been questioned in a study by O'Connell (1997) who showed that the standard practice of calculating real exchange rates relative to the US dollar can lead to cross-sectional dependence in time series panel data.

Table 3
Cross-sectional tests of PPP

	Horizon (quarters)							
	1		4		8		12	
Country	$c = 0$	$c \neq 0$	$c = 0$	$c \neq 0$	$c = 0$	$c \neq 0$	$c = 0$	$c \neq 0$
United States	−0.64	−1.11	−0.24	−0.92	−0.16	−0.92	−0.17	−0.98
	(0.35)	(0.61)	(0.37)	(0.49)	(0.41)	(0.54)	(0.39)	(0.50)
Australia	−0.22	0.20	−0.05	−0.04	0.00	−0.33	0.01	−0.42
	(0.29)	(0.53)	(0.32)	(0.41)	(0.32)	(0.32)	(0.28)	(0.31)
Austria	−0.98	−0.87	−0.50	−0.33	−0.38	−0.27	−0.37	−0.38
	(0.25)	(0.33)	(0.34)	(0.51)	(0.28)	(0.47)	(0.26)	(0.36)
Belgium	−0.68	−0.67	−0.21	−0.12	−0.10	−0.03	−0.09	−0.06
	(0.24)	(0.31)	(0.27)	(0.37)	(0.25)	(0.36)	(0.26)	(0.35)
Canada	−0.69	−1.78	−0.15	−1.11	−0.07	−1.10	−0.06	−1.11
	(0.36)	(0.61)	(0.41)	(0.53)	(0.44)	(0.51)	(0.40)	(0.49)
Denmark	−0.68	−0.73	−0.24	−0.14	−0.15	−0.07	−0.12	−0.07
	(0.16)	(0.18)	(0.19)	(0.22)	(0.17)	(0.23)	(0.15)	(0.21)
Finland	−0.68	−1.00	−0.33	−0.67	−0.25	−0.63	−0.21	−0.65
	(0.16)	(0.25)	(0.23)	(0.33)	(0.21)	(0.37)	(0.16)	(0.31)
France	−0.44	−0.23	−0.13	−0.01	−0.06	−0.02	−0.06	−0.05
	(0.20)	(0.28)	(0.23)	(0.26)	(0.21)	(0.23)	(0.18)	(0.19)
Germany	−0.98	−0.88	−0.36	−0.21	−0.16	0.07	−0.05	0.01
	(0.34)	(0.49)	(0.46)	(0.67)	(0.39)	(0.55)	(0.38)	(0.40)
Italy	−0.33	−0.36	−0.10	−0.20	−0.05	−0.21	−0.05	−0.23
	(0.16)	(0.25)	(0.15)	(0.27)	(0.15)	(0.28)	(0.13)	(0.27)
Japan	−0.83	−0.58	−0.43	−0.18	−0.37	−0.08	−0.43	−0.08
	(0.25)	(0.28)	(0.24)	(0.26)	(0.23)	(0.20)	(0.22)	(0.18)
Netherlands	−0.89	−0.88	−0.35	−0.36	−0.23	−0.32	−0.23	−0.40
	(0.26)	(0.31)	(0.33)	(0.37)	(0.28)	(0.33)	(0.26)	(0.24)
Norway	−0.68	−1.03	−0.27	−0.58	−0.18	−0.63	−0.16	−0.64
	(0.20)	(0.24)	(0.25)	(0.24)	(0.23)	(0.23)	(0.18)	(0.17)
Spain	−0.43	−0.44	−0.20	−0.17	−0.15	−0.21	−0.15	−0.25
	(0.22)	(0.43)	(0.19)	(0.27)	(0.16)	(0.27)	(0.13)	(0.24)
Sweden	−0.47	−0.66	−0.08	−0.21	−0.02	−0.48	0.00	−0.45
	(0.23)	(0.31)	(0.26)	(0.45)	(0.23)	(0.50)	(0.18)	(0.43)
Switzerland	−1.17	−0.79	−0.77	−0.54	−0.63	−0.40	−0.60	−0.29
	(0.36)	(0.54)	(0.31)	(0.52)	(0.29)	(0.55)	(0.29)	(0.55)
United Kingdom	−0.59	−0.89	−0.26	−0.63	−0.18	−0.52	−0.18	−0.51
	(0.17)	(0.27)	(0.23)	(0.28)	(0.25)	(0.34)	(0.24)	(0.34)
Pooled	−0.43	−0.64	−0.10	−0.33	−0.04	−0.31	−0.03	−0.32
	(0.15)	(0.16)	(0.16)	(0.19)	(0.16)	(0.21)	(0.13)	(0.21)

Sample period is 1972.I–1996.III. Entries report numeraire invariant regression coefficients β_i of the effect of inflation on the real exchange rate in the model $\Delta_k q_{ij} = c_i - c_j + \beta_i \Delta_k p_i - \Delta_k p_j + \nu_{ij}$. Price index is the CPI. The column headings ($c \neq 0$) or ($c = 0$) indicate whether or not the regression contains a constant term (normalized using $c_0 = 0$). Standard errors have been computed using the Newey–West procedure and are reported in parentheses.

In this paper we investigate PPP among 17 currencies between 1972 and 1996 using a panel framework that explicitly deals with the numeraire effect that causes the cross-sectional dependence. Our results are independent of the specific characteristics of the chosen benchmark currency. We find that for a system of multiple exchange rates PPP provides a relatively accurate description of exchange rate movements. This hold especially at horizons longer than 1 year. The extent to which this holds differs, however, from currency to currency or from currency-blocs to currency-blocs. In other words, we find that the choice of currencies is crucial in PPP research.

Our evidence in favor of PPP is strongest for the German mark, and much weaker for the US dollar. Lothian (1998) has recently suggested that the difficulty of finding evidence of PPP with the United States dollar as the numeraire currency is caused by the 1980–1987 period during which the dollar first strongly depreciated and afterwards strongly appreciated. The least evidence in favor of PPP is found when the Japanese yen is used as numeraire. In addition to PPP, one needs a trend to explain movements in this currency.

On the basis of our findings we conclude that there is substantive evidence that PPP holds for many currencies, although not for every currency to the same extent. Instead of concentrating on its general validity, research on PPP should now try to explain why it holds within currency blocs and not between them. Potential explanations are the fact that goods arbitrage is more effective between for instance European countries due to their proximity and the lower volatility of these currencies. We intend to investigate this issue in future research.

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