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Adjusting toward long-run purchasing power parity

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A B S T R A C T

Under purchasing power parity (PPP) exchange rates and relative prices adjust to maintain a constant real exchange rate in the long run. Its empirical validity continues to be questioned. We use data on exchange rates and prices relative to the U.S. for a long-span (1870–2020) panel of 16 countries to examine (a) whether the long-run elasticity is one; (b) whether there is adjustment by exchange rates or prices to maintain a constant real exchange rate and (c) the time taken to adjust. We use four estimators, which increasingly restrict the model. These are country-specific vector error correction model in exchange rates and relative prices; the Johansen estimator, which has the cross-equation restriction that the long-run coefficient in the two equations is the same; the system pooled mean group estimator, which has a homogeneous long-run coefficient over countries and heterogeneous short-run dynamics, and a univariate real exchange rate equation used to obtain median unbiased estimates of the half-life.

1. Introduction

Relative purchasing power parity (PPP) suggests that exchange rates and relative prices between countries adjust so that in the long run the real exchange rate is constant. Vo and Vo (2022, p.20) state that "about one-half the total citations count among prominent exchange rate research are attributable to studies dedicated to testing the empirical validity of PPP." But its empirical validity continues to be questioned (Burstein and Gopinath, 2014; Itskhoki and Mukhin, 2021; Rogoff, 1996), spawning a body of literature as to why PPP does not hold, at least in short to medium-run (Balassa, 1964; Harding et al., 2020; Samuelson, 1964).

Because of the persistent nature of the data, tests for unit roots in the real exchange rate and for cointegration between exchange rates and relative prices lack power, (Engel and Morley, 2001; Imbs et al., 2005; Taylor and Taylor, 2004). We increase power by using long-span panel data for consumer price index differentials and exchange rates, relative to the US, for 16 advanced economies over 151 years (1870–2020), taken from the Jordà-Schularick-Taylor macrohistory database (Jordà et al., 2017).

Purchasing power parity is a composite hypothesis. One part is whether the long-run elasticity between exchange rates and relative prices equals one, the other part is whether, given a unit long-run elasticity, relative prices or exchange rates adjust to keep the real exchange rates constant. If they do adjust we are also interested in the time taken for the real exchange rate to adjust.

We answer these questions in four stages. Using increasingly restricted models. In the most general model, we estimate an unrestricted bivariate vector error correction model for each country. The price and exchange rate equations in this system give different long-run elasticities. However, these are similar on average, and we impose the cross-equation restriction that they are equal using the Johansen procedure. The estimates of the long-run elasticity average close to one over the 16 countries but are quite dispersed, which

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is not surprising given the eventful history over this period, which includes wars, depressions and many types of exchange rate regimes.

We then pool the long-run estimates while leaving the short-run dynamics unrestricted using the System Pooled Mean Group (SPMG) estimator recently proposed by Chudik et al. (2023). This estimates a homogeneous long-run relationship between two endogenous variables while permitting heterogeneous short-run dynamics across countries. The more efficient, pooled long-run elasticity between exchange rates and relative prices is precisely estimated and not significantly different from one: 0.990 with a 95% bootstrapped confidence interval of 0.978–1.002. The bootstrapped confidence interval allows cross-section dependence, which is evident, perhaps because of a dollar factor.

Papell and Prodan (2020) use similar data (1870–2013) to us but a different estimator: the covariability approach of Müller and Watson (2018). They find that the unit long-run elasticity between exchange rate depreciation and inflation differential cannot be rejected for 9 of 16 countries at a 10% significance level.

SPMG allows episodic failures when cointegration between exchange rates and relative prices does not hold. Chudik et al. (2023, p.1024) note that episodic failures include “major shocks such as wars, depressions, natural disasters, or important policy failures.”.

Since SPMG permits heterogeneous short-run dynamics and two-way causality, we can examine whether relative prices or exchange rates adjust in each country. Relative prices adjust to maintain a constant real exchange rate in all 16 countries, and only Germany and Italy do not show significant exchange rate adjustments. This could be because exchange rates are pegged around half of the sample period in most countries.[1] Even if the exchange rate is destabilising, the system can be stable because of price adjustments.

We examine the changing pattern of adjustment using a thirty-year window rolling estimator. In Germany and Italy, the exchange rates adjusted to the parity during the floating exchange rate regime. The evidence for adjusting toward relative PPP strengthens once we consider exchange rate regime.

To examine the time taken to adjust, we combine the exchange rate and price differential equations to give a real exchange rate equation that allows us to obtain a median unbiased estimate of the half-life: the number of years for a shock to the real exchange rate to dissipate by half. This is not well defined when two separate equations with two separate shocks are estimated, as was done above. Long-span data is particularly valuable for estimating the half-life. Using quarterly post-1973 real exchange rates, Murray and Papell (2002) and Rossi (2005) find that the upper confidence limit of the half-life is infinite. Steinsson (2008, p.521) says that “even 30 years after the breakdown of Bretton Woods, it is not possible to estimate the half-life of the real exchange rate with much precision.” Using the median unbiased estimates of Hansen (1999), we find finite half-lives for 11 of 16 countries based on 95% confidence intervals. Our median 95% lower bound, 3.16 years, is consistent with Lopez et al. (2013), who cover the same countries (1870–1998). But we can reject unit root in more countries; the margin of error and upper confidence limit are also much lower. This is owing to a larger span of data, that is particularly valuable if we want to estimate half-life with some precision.

The rest of the paper is organised as follows. Section 2 describes the data. Section 3 sets out the models. Section 4 presents the estimates. Section 5 concludes.

2. Data

We use spot exchange rates and consumer price indices from the Jordà-Schularick-Taylor macrohistory database. Using the U.S. as the benchmark, an almost balanced panel[2] comprises N=16 and T=151 (1870–2020). Ireland has a shorter span (T=95), therefore, we exclude it from the sample. The exchange rate is quoted in local currency per U.S. dollar; thus, an increase in the exchange rate represents a depreciation against the dollar. Prices are measured with Consumer Price Indexes (1990 = 100). Given the data, we can only test for relative purchasing power parity that considers changes in absolute parity:

$$\frac{\Delta E}{E} = \frac{\Delta P}{P} - \frac{\Delta P^*}{P^*}$$

The percentage change in expected exchange rates would mirror the inflation differential between two countries. If the inflation rate is higher in the UK than in the US, the pound is expected to depreciate against the US dollar.

Figure A1 plots the exchange rate depreciation and inflation differentials across 16 countries to the U.S. Table A1 shows the summary statistics for the proportionate changes in exchange rates and inflation differentials. Wars dominate large changes in exchange rates and inflations.

The variables used for estimation are the logarithm of the spot exchange rate of country *i* against the US dollar, and the difference of the logarithm of the price indices where the foreign price index is the U.S. CPI. Tests for unit roots and cointegration have low power and panel unit root tests can only reveal whether all are I(1), the null, or some proportion are I(0). We proceed on the assumption that both these series are I(1). The log price differential will be I(1) either if both are I(1) or they are I(2) and cointegrate to I(1). If the exchange rate and price differential cointegrate, there must be at least one non-zero adjustment coefficient, which we will investigate.

[1] the gold standard (before 1940), the Bretton Woods system (1951–1971), the Exchange Rate Mechanism (1979–1988) and the eurozone (1999-).

[2] 2413/2416 exchange rate observations. We interpolate the 1945 exchange rate value using the average of 1944 and 1946 values. The missing observations are the 1871 and 1872 values of exchange rates in Japan.

3. Models

We use four estimators, which increasingly restrict the model. These are the country-specific Vector Error Correction Model (VECM); the Johansen VECM; the System Pooled Mean Group estimator; and a univariate real exchange rate equation.

3.1. The vector error correction model (VECM)

Consider y_{it} as the log of the spot exchange rate, and x_{it} as the log price differential ($\log p_{it} - \log p^{*}_{it}$) where p^{*}_{it} is the U.S. CPI, the VECM of exchange rate and relative prices are:

$$\Delta y_{it} = a_{10,i} + a_{11,i}\left(y_{i,t-1} - \theta_{yi}x_{i,t-1}\right) + \sum_{\ell=1}^{p-1}\psi'_{1\ell i}\Delta w_{i,t-\ell} + u_{1it} \tag{1}$$

$$\Delta x_{it} = a_{20,i} + a_{21,i}\left(y_{i,t-1} - \theta_{xi}x_{i,t-1}\right) + \sum_{\ell=1}^{p-1}\psi'_{2\ell i}\Delta w_{i,t-\ell} + u_{2it} \tag{2}$$

where $\Delta y_{it} = y_{it} - y_{i,t-1}$, $\Delta x_{it} = x_{it} - x_{i,t-1}$ and $\Delta w_{it} = (\Delta y_{it}, \Delta x_{it})'$. Both equations have a constant ($a_{10,i}, a_{20,i}$), which is important to test for relative PPP because our data are price indices. This system's exchange rate and price equations give different long-run elasticities θ_{yi} and θ_{xi}, whereas PPP implies a common elasticity.

3.2. Johansen VECM

The cross-equation restriction that the long-run elasticity is equal across the exchange rate and relative prices equations, $\theta_{yi} = \theta_{xi} = \theta_i$, can be imposed using the Johansen (1991) procedure:

$$\Delta y_{it} = a_{10,i} + a_{11,i}\left(y_{i,t-1} - \theta_i x_{i,t-1}\right) + \sum_{\ell=1}^{p-1}\psi'_{1\ell i}\Delta w_{i,t-\ell} + u_{1it}$$

$$\Delta x_{it} = a_{20,i} + a_{21,i}\left(y_{i,t-1} - \theta_i x_{i,t-1}\right) + \sum_{\ell=1}^{p-1}\psi'_{2\ell i}\Delta w_{i,t-\ell} + u_{2it} \tag{3}$$

The PPP null hypothesis is $H_0 : \theta_i = 1$, a unit long-run elasticity.

3.3. The System Pooled Mean Group (SPMG) estimator

The next step pools the long-run estimates while leaving the short-run dynamics unrestricted using the System Pooled Mean Group (SPMG) estimator recently proposed by Chudik et al. (2023). This estimates a homogeneous long-run relationship, θ, between two endogenous variables while permitting heterogeneous short-run dynamics across countries:

$$\Delta y_{it} = a_{10,i} + a_{11,i}\left(y_{i,t-1} - \theta x_{i,t-1}\right) + \sum_{\ell=1}^{p-1}\psi'_{1\ell i}\Delta w_{i,t-\ell} + u_{1it}$$

$$\Delta x_{it} = a_{20,i} + a_{21,i}\left(y_{i,t-1} - \theta x_{i,t-1}\right) + \sum_{\ell=1}^{p-1}\psi'_{2\ell i}\Delta w_{i,t-\ell} + u_{2it} \tag{4}$$

The SPMG model allows testing of $H_0 : \theta_i = \theta = 1$ for all i.

The SPMG model allows two-way long-run causality between exchange rates and relative prices. Exchange rates or prices can adjust to parity. Auer et al. (2021) dissect the adjustment of import and retail prices following a large and sudden appreciation of the Swiss franc in January 2015. This is consistent with relative prices adjusting to exchange rate changes.

If we do not reject the unit long-run elasticity between exchange rates and relative prices from the SPMG model, we could impose a unit coefficient on the exchange rate and relative prices ($\theta = 1$) in equation (4):

$$\Delta y_{it} = a_{10,i} + a_{11,i}\left(y_{i,t-1} - x_{i,t-1}\right) + \sum_{\ell=1}^{p-1}\psi'_{1\ell i}\Delta w_{i,t-\ell} + u_{1it}$$

$$\Delta x_{it} = a_{20,i} + a_{21,i}\left(y_{i,t-1} - x_{i,t-1}\right) + \sum_{\ell=1}^{p-1}\psi'_{2\ell i}\Delta w_{i,t-\ell} + u_{2it} \tag{5}$$

We test relative purchasing power parity by looking at exchange rate changes (Δy_{it}) and inflation differential (Δx_{it}) adjusting to the parity ($y_{i,t-1} - x_{i,t-1}$). Here, the object is to examine the relative extent to which exchange rates or relative prices adjust to maintain

Table 1
Country=Specific Elasticities from the VECM.

	Unrestricted VECM				Restricted VECM	
	Exchange rate equation (1)		Inflation equation (2)		Johansen model (3)	
country	LR. Elast.	Std errors	LR. Elast.	Std errors	LR. Elast.	90% CI
Australia	0.941	0.240	1.127	0.157	1.072	0.873, 1.272
Belgium	0.710	0.124	0.928	0.089	0.850	0.763, 0.937
Canada	0.964	0.305	3.621	2.596	1.396	0.909, 1.883
Denmark	0.462	0.146	0.831	0.417	0.531	0.326, 0.735
Finland	0.964	0.027	2.076	9.103	0.970	0.928, 1.013
France	1.001	0.013	0.942	0.083	1.005	0.984, 1.025
Germany	1.004	0.008	1.002	0.007	0.995	0.987, 1.003
Italy	0.951	0.035	0.981	0.022	0.971	0.945, 0.996
Japan	0.782	0.058	0.980	0.171	0.814	0.738, 0.890
Netherlands	−0.020	0.613	1.786	1.255	0.591	−0.202, 1.384
Norway	0.594	0.139	1.239	0.950	0.645	0.427, 0.863
Portugal	0.942	0.068	1.131	0.236	0.979	0.901, 1.057
Spain	0.950	0.102	0.981	0.082	0.968	0.875, 1.061
Sweden	0.706	0.098	1.271	0.742	0.754	0.600, 0.908
Switzerland	3.429	2.385	2.182	0.293	2.249	1.784, 2.714
UK	0.904	0.057	0.940	0.124	0.912	0.833, 0.991
Average	0.955		1.376		0.981	
Std. error	0.178		0.182		0.099	

Notes: This table shows the long-run elasticity estimates from the unrestricted VECM (equations (1) and (2)) and the Johansen model (equation (3)).

PPP. The heterogeneous adjustment coefficients allow some countries to adjust and others not to.

3.4. Univariate real exchange rate equation

We can combine the exchange rates and prices equations in (5) to give a real exchange rate equation to obtain a median unbiased estimate of the half-life.[3] Consider $q_{it} = y_{i,t} - x_{i,t}$ as the log of real exchange rate, the univariate equation is:

$$\Delta q_{it} = b_{0,i} + b_{1,i} q_{i,t-1} + \sum_{\ell=1}^{p-1} \varphi_{\ell i} \Delta q_{i,t-\ell} + u_{it} \tag{6}$$

where the adjustment is the parameter of interest. Like equations (5), it imposes unit long-run elasticity between relative prices and exchange rates. The null hypothesis is $b_{1,i} = 0$ and the alternative is $b_{1,i} < 0$. If the real exchange rate mean reverts, then there is evidence of adjusting toward relative PPP. Equation (6) is also an Augmented Dickey-Fuller (ADF) regression of the real exchange rate and the stage two test in Froot and Rogoff (1995).

We report the median unbiased estimates and their half-lives, defined as the number of years for a shock to dissipate by half. We calculate the median unbiased estimates of $b_{0,i}$, $b_{1,i}$ and $\varphi_{\ell i}$ in equation (6) using the grid-bootstrap method described in Hansen (1999), which is closely related to the method proposed by Andrews and Chen (1994).

4. Results and discussion

4.1. Country-Specific long-run elasticities

Table 1 shows that using unrestricted VECM, the country-specific long-run elasticities in the two equations are quite dispersed.[4] This is common in cross-country studies and the outliers tend to have large standard errors. The average long-run elasticity from the exchange rate equation is 0.955 (standard error 0.178), which is not significantly different from one. The average from the inflation equation is 1.376 (standard error 0.182) is just not significantly different from one using a t distribution with 15 degrees of freedom. The cross-equation restriction that the long-run elasticities from the two equations are the same is imposed through the use of the Johansen procedure. The cross-country average of the Johansen estimates is 0.981 (0.099). Again it is not significantly different from one, but more precisely estimated with a smaller cross-country dispersion than in the individual equations. However, considerable dispersion remains and the 90% confidence interval does not cover a unit elasticity in 7 out of 16 countries. The large shocks and episodes where cointegration fails could be an explanation and pooling the data may reduce these problems. We, therefore, use the SPMG estimator to estimate a common long-run elasticity across countries.

[3] The dimension reduction that is involved in going from equation (5) to equation (6) is discussed in Boyd and Smith (1999, p.291–292). The real exchange rate equation imposes common speeds of adjustment in the SPMG model.

[4] We use lag length two to estimate the long-run elasticities in the unrestricted VECM (equations (1) and (2) and Johansen VECM (equation (3).

Table 2
Speed of adjustment of spot exchange rates and inflation differential to relative PPP.

		Exchange rate			Inflation differential		
country	lag	speed	90% CI	t.	speed	90% CI	t.
Australia	0	−0.064	−0.124, −0.004	−1.77	0.053	0.025, 0.081	3.13
Belgium	1	−0.046	−0.088, −0.003	−1.77	0.082	0.042, 0.122	3.37
Canada	0	−0.144	−0.216, −0.072	−3.31	0.046	0.011, 0.080	2.16
Denmark	6	−0.019	−0.068, 0.030	−0.63	0.019	0.001, 0.037	1.70
Finland	3	−0.177	−0.265, −0.090	−3.35	0.019	−0.043, 0.081	0.50
France	0	−0.177	−0.294, −0.060	−2.49	0.110	0.069, 0.151	4.41
Germany	5	0.758	0.062, 1.454	1.79	0.970	0.250, 1.690	2.22
Italy	0	0.048	−0.046, 0.143	0.84	0.303	0.261, 0.345	11.93
Japan	2	−0.019	−0.051, 0.013	−0.96	0.046	−0.001, 0.094	1.60
Netherlands	1	−0.083	−0.134, −0.033	−2.73	0.019	0.002, 0.036	1.87
Norway	6	−0.058	−0.119, 0.002	−1.58	0.014	−0.009, 0.037	0.98
Portugal	1	−0.091	−0.156, −0.026	−2.32	0.018	−0.025, 0.061	0.68
Spain	7	−0.039	−0.122, 0.045	−0.76	0.048	0.021, 0.075	2.92
Sweden	1	−0.142	−0.205, −0.079	−3.70	0.029	0.004, 0.055	1.87
Switzerland	1	−0.005	−0.034, 0.025	−0.26	0.018	0.004, 0.031	2.15
UK	8	−0.158	−0.280, −0.036	−2.13	0.039	−0.001, 0.078	1.62

Notes: This table shows the speed of adjustment, ninety percent confidence intervals and t-statistics of exchange rates and relative prices adjusting to PPP deviations from equation (5). Tables A3 and A4 in the Appendix show the full estimation results.

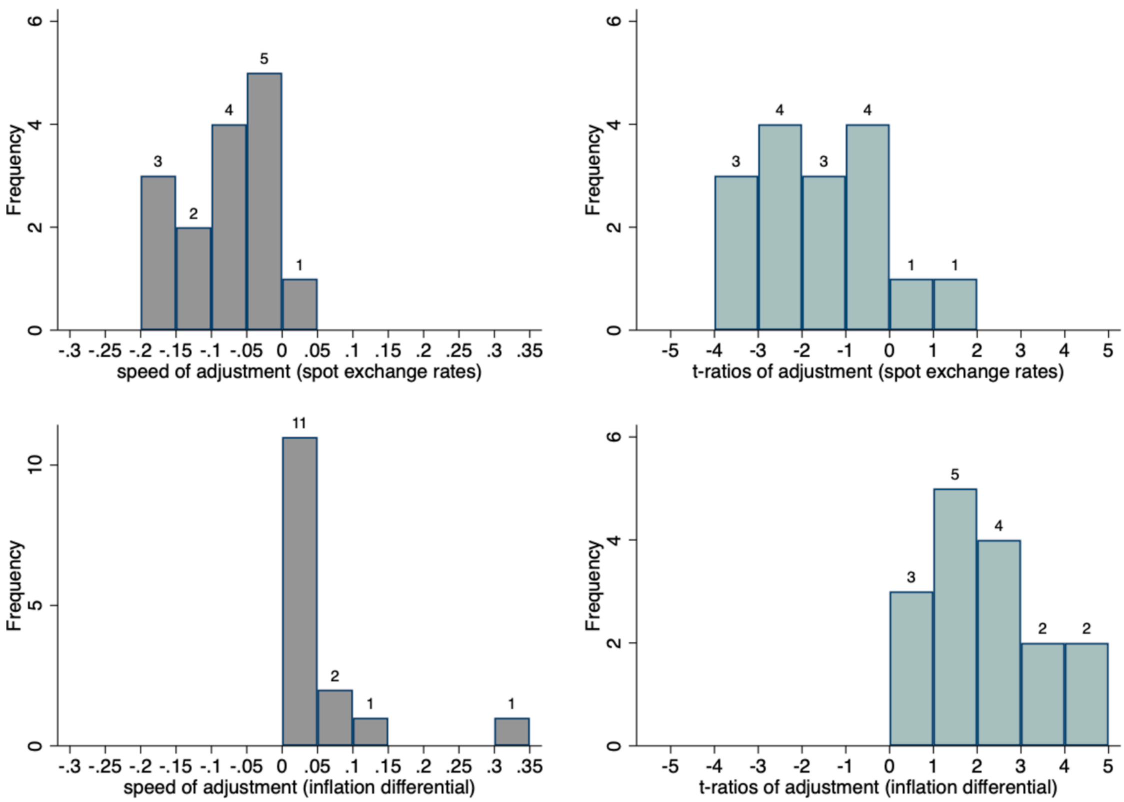

Fig. 1. Histogram of *t*-ratios and speed of adjustment of spot exchange rates and inflation differential to relative PPP. Notes: Fig. 1 presents the distribution of speeds and t-ratios of adjustment to PPP deviations, the parameters of interest, from Table 2. We exclude Germany's exchange rate and inflation differential adjustments from the speed of adjustment because their estimates are large (0.758 and 0.970, respectively). The t-stats of Italy's inflation differential adjustment are capped at 5 for visibility as it takes the value of 11.93.

Table 3
Speed of adjustment of real exchange rate to relative PPP.

country	lag	speed	90% CI	t.
Australia	0	−0.117	−0.053, −0.181	−3.03
Belgium	1	−0.134	−0.082, −0.186	−4.25
Canada	0	−0.190	−0.112, −0.268	−4.00
Denmark	6	−0.045	0.006, −0.096	−1.46
Finland	3	−0.208	−0.108, −0.309	−3.41
France	0	−0.287	−0.192, −0.382	−4.99
Germany	5	−0.201	−0.109, −0.293	−3.59
Italy	0	−0.255	−0.165, −0.345	−4.66
Japan	2	−0.053	0.007, −0.113	−1.44
Netherlands	1	−0.098	−0.047, −0.149	−3.14
Norway	6	−0.077	−0.013, −0.142	−1.98
Portugal	1	−0.105	−0.047, −0.163	−2.99
Spain	7	−0.093	−0.013, −0.173	−1.91
Sweden	1	−0.172	−0.106, −0.239	−4.25
Switzerland	1	−0.030	0.001, −0.062	−1.58
UK	8	−0.234	−0.112, −0.356	−3.16

Notes: This table shows the speed of adjustment of real exchange rates on equation (6).

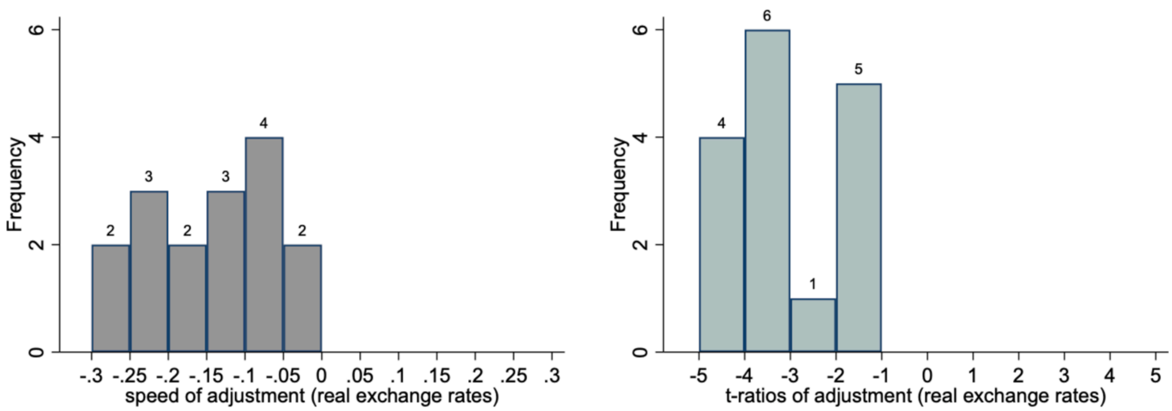

Fig. 2. Histogram of *t*-ratios and speed of adjustment of real exchange rates to relative PPP. Notes: Fig. 2 presents the distribution of speeds and t-ratios of adjustment to PPP deviations, the parameters of interest, from Table 4.

4.2. The SPMG estimator

The SPMG long-run elasticity point estimate of equation (4) is 0.990.[5] PPP is always normalised on the exchange rate. Therefore, we use the direct estimate of the long-run coefficient (y on x) from Table A2. There are significant degrees of cross-section dependence; therefore, we focus on the bootstrapped 95% confidence interval. The bootstrapped 95% confidence interval is tight (0.978–1.002). The SPMG long-run elasticity estimate supports the relative PPP, with a tight range that is very close to 1.

This gives us the confidence to impose a unit long-run elasticity across countries and examine the evidence of relative PPP based on the adjustments. Imposing a unit long-run elasticity also increases efficiency. We use a general-to-specific approach to select the lag length for each country in equation (5). We keep the lag length the same between the SPMG and the real exchange rate equation (6) to compare their speed of adjustments. The SPMG model selects a longer lag than the univariate model. This is because the real exchange rate fluctuates less than its two components (Figures A1 vs. A2). Therefore, we use the lag length of the univariate model. We use the 5% level of significance. The lags vary between 0 and 8 (Table 2).

Three main findings can be drawn. First, we find evidence of relative PPP in 16 countries through either exchange rates or relative prices adjusting to the parity (see Table 2). Second, the evidence of adjusting toward parity is more common from the relative price than the exchange rate channel. All countries' relative prices adjust to PPP, but the exchange rates in Germany and Italy did not adjust to PPP. On statistical significance, 8 countries' relative price channel lap the threshold of significance, two, for instance, whilst 7 do in the exchange rate channel (see Fig. 1). Third, the exchange rate does the bulk of adjustment to the parity. 9 countries' exchange rate adjusts faster than 5%, but 11 countries' inflation adjusts slower than 5% (Fig. 1). Friedman (1953) argues that internal prices adjust slower than exchange rates when arguing for flexible exchange rates.

[5] Following Chudik et al. (2023), the SPMG model uses one lag in equation (4).

Table 4
Median unbiased half-lives of real exchange rates.

$$q_t = c + \alpha q_{t-1} + \sum_{i=1}^{k} \psi_i \Delta q_{t-i} + u_t$$

	Country	Lag	α_{OLS}	α_{MU}	95% CI	$HL_{IRF,MU}$	95% CI
1	France	0	0.713	0.732	[0.643, 0.824]	2.25	[1.62, 3.60]
2	Italy	0	0.745	0.763	[0.675, 0.855]	2.59	[1.80, 4.45]
3	Finland	3	0.792	0.818	[0.722, 0.916]	2.63	[2.02, 6.68]
4	Germany	5	0.799	0.820	[0.729, 0.907]	3.69	[3.16, 8.64]
5	UK	8	0.766	0.798	[0.696, 0.918]	3.69	[1.86, 9.65]
6	Canada	0	0.810	0.829	[0.740, 0.913]	3.71	[2.34, 7.66]
7	Sweden	1	0.828	0.843	[0.758, 0.914]	4.74	[3.22, 8.36]
8	Belgium	1	0.866	0.877	[0.800, 0.928]	6.23	[4.21, 10.12]
9	Australia	0	0.883	0.904	[0.813, 0.972]	6.86	[3.37, 24.60]
10	Portugal	1	0.895	0.914	[0.825, 0.974]	8.32	[4.08, 26.84]
11	Netherlands	1	0.902	0.918	[0.832, 0.969]	8.92	[4.63, 23.56]
12	Spain	7	0.907	0.955	[0.837, 1.000]	12.18	[2.86, ∞]
13	Norway	6	0.923	0.959	[0.853, 1.000]	14.60	[4.55, ∞]
14	Switzerland	1	0.970	0.994	[0.919, 1.000]	121.71	[9.07, ∞]
15	Denmark	6	0.955	0.998	[0.885, 1.000]	∞	[5.05, ∞]
16	Japan	2	0.947	0.999	[0.877, 1.000]	∞	[0.89, ∞]

Notes: The half-lives are displayed in ascending order. The median unbiased estimates of the ADF regression are computed using the grid-bootstrap method described in Hansen (1999). α_{OLS} is the least squares estimate, α_{MU} is the median-unbiased estimate, $HL_{IRF,MU}$ is the half-life estimate from the impulse response function. These half-lives were generated using Steinsson's (2008) programs.

We find evidence of PPP in more countries than in Papell and Prodan (2020). Considering both adjustment channels, we find evidence of PPP in 14 out of 16 countries at a 10% significance level. When the confidence interval of adjustment contains zero, we are unsure about the evidence for relative PPP. We find evidence of PPP in 9 out of 16 countries based on exchange rate adjustment and in 11 out of 16 countries based on relative price adjustment. We are unsure if there is an adjustment to PPP in Japan and Norway. In contrast, Papell and Prodan (2020) find evidence of PPP in Japan. Our paper complements Papell and Prodan (2020) using a different approach on the same dataset.

4.3. The univariate real exchange rate equation

We learn two things by comparing the SPMG model (5) estimates with the univariate real exchange rate equation (6). First, we distinguish the speed of adjustment by prices and exchange rates.[6] In Germany and Italy, relative prices drive the real exchange rate adjustment to parity. Even if one channel is destabilising, as with Germany's and Italy's exchange rate, the net effect can stabilise if the other adjusts fast enough. Therefore, the real exchange rate in 16/16 countries adjust to PPP (Table 3). Second, we can be more confident about the evidence of relative PPP based on the univariate approach. 11 and 10 countries lap the thresholds, two and three, respectively (Fig. 2). However, when we are unsure about a real exchange rate adjustment in Denmark, Japan, and Switzerland (the 90% confidence interval contains zero), we can be more confident based on the relative prices adjusting to PPP in these countries.

Table 4 reports the median unbiased estimates and half-lives for the real exchange rate in equation (6).[7] The least squares estimates are biased downward relative to the median unbiased estimates. Using a 95% confidence interval, we can't be sure that the half-life is finite in 5/16 countries: Denmark, Japan, Norway, Spain and Switzerland. These are also the countries where we do not reject the unit root null hypothesis based on the ADF critical values (Table 3). They also have the longest half-lives (from 12.11 years in Norway to infinity in Japan). For the other eleven countries where we are confident that the half-life is finite,[8] the point estimates range from 2.25 years (France) to 8.92 years (Netherlands). The average is 4.87 years, which is within the "consensus range" of 3 to 5 years (Rogoff, 1996). The margins of error (95% confidence intervals) are less than 3 years in 8/11 countries and around 10 years in 3/11 countries.

Lopez et al. (2013) is our closest comparison because it examines the same countries with long-span data (1870–1998) and uses consumer price indices. They reject the unit root in 9/16 countries (8/9 countries overlap with ours); we reject the unit root in 11/16 countries. Our lower confidence limit (95%) ranges between 1.62 and 4.63 years. The median lower confidence limit is 3.16 years, which is similar to Lopez et al. (2013). Our upper confidence limit (95%) ranges between 3.60 and 26.84 years, with a median of 8.64 years. In contrast, their upper confidence limit ranges between 5.69 and 72 years, with a median of 18.22 years.[9] Our margin of error

[6] Because the SPMG and the univariate approaches have the same variables, the coefficients (not the t-ratios) will add up. To illustrate, the real exchange rate adjustment speed (−0.117) sums up the adjustment speed of the exchange rate (−0.064) and inflation (0.053) in Australia. But they may differ because of the restrictions discussed in Pesaran and Smith (1998, p.496–8). Rounding matters also matter with small numbers like this.

[7] Half-lives are not well-defined in multivariate (SPMG) models; therefore, we only look at the half-life of the real exchange rate. It does not make sense to look at the half-life of the exchange rate and relative prices because both are I(1), so their half-lives are infinite.

[8] We see evidence of mean reversion in countries with finite half-lives on the real exchange rate plots (Figure A2).

[9] Table 4. Median-unbiased half-lives in ADF regressions of Lopez et al. (2013).

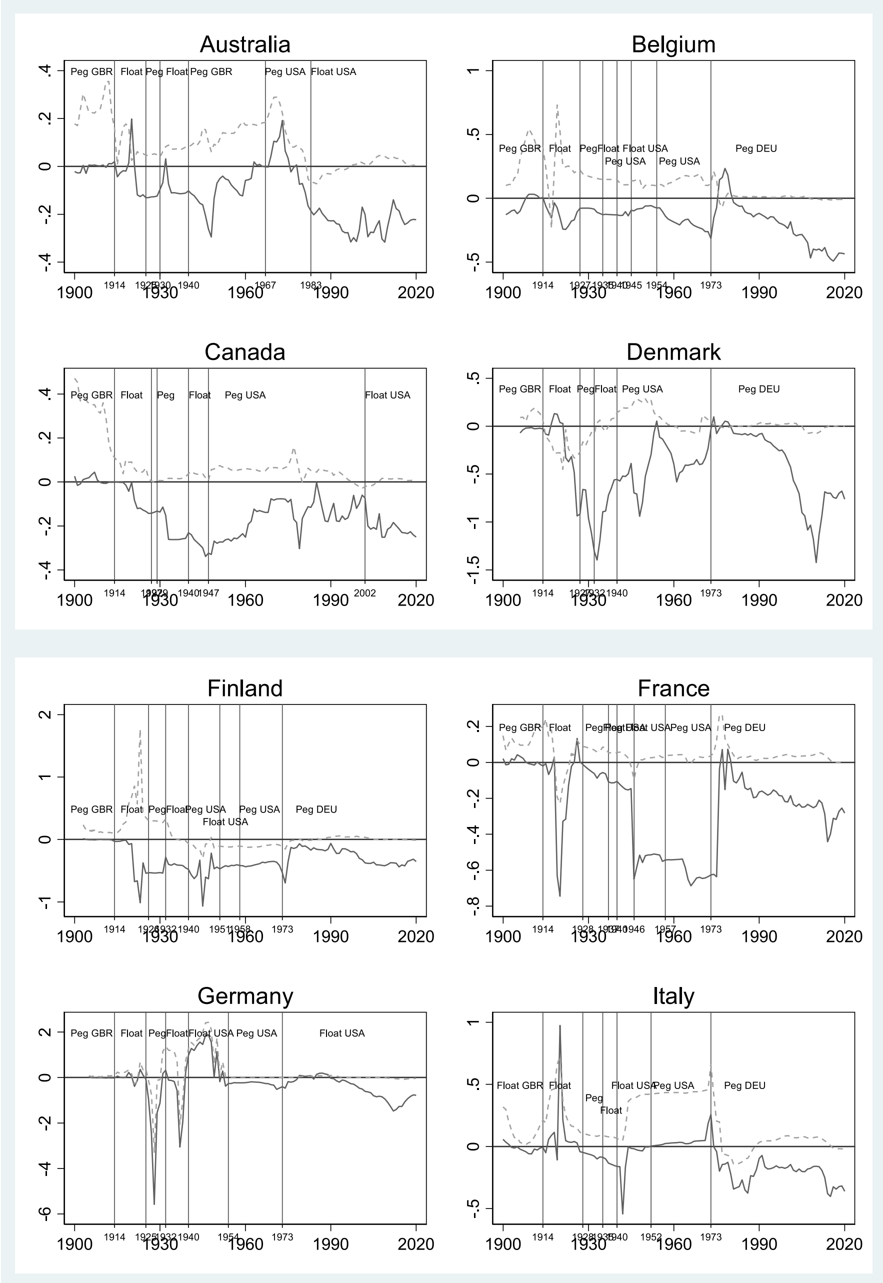

Fig. 3. Rolling adjustment of spot exchange rates and inflation differential to deviation from PPP. Notes: The solid line indicates exchange rates, and the short-dashed line indicates relative prices. The exchange rate regimes and the base currencies are labelled following the Jordà-Schularick-Taylor macro-financial dataset (Jordà et al., 2017).

and upper confidence limit are much lower than theirs. This is likely because the span of our data is larger: 151 years (1870–2020) vs. 129 years (1870–1998) in their study.

4.4. Rolling adjustments — Subsamples and an alternative estimator

One plausible reason we find more common evidence of adjustment via relative prices than exchange rates could be that the exchange rate is pegged half the time in many countries (see Table A5). The pegged exchange rate includes the gold standard (before 1940), the Bretton Woods system (1951–1971), the Exchange Rate Mechanism (1979–1998) and the eurozone (1999-). A pegged

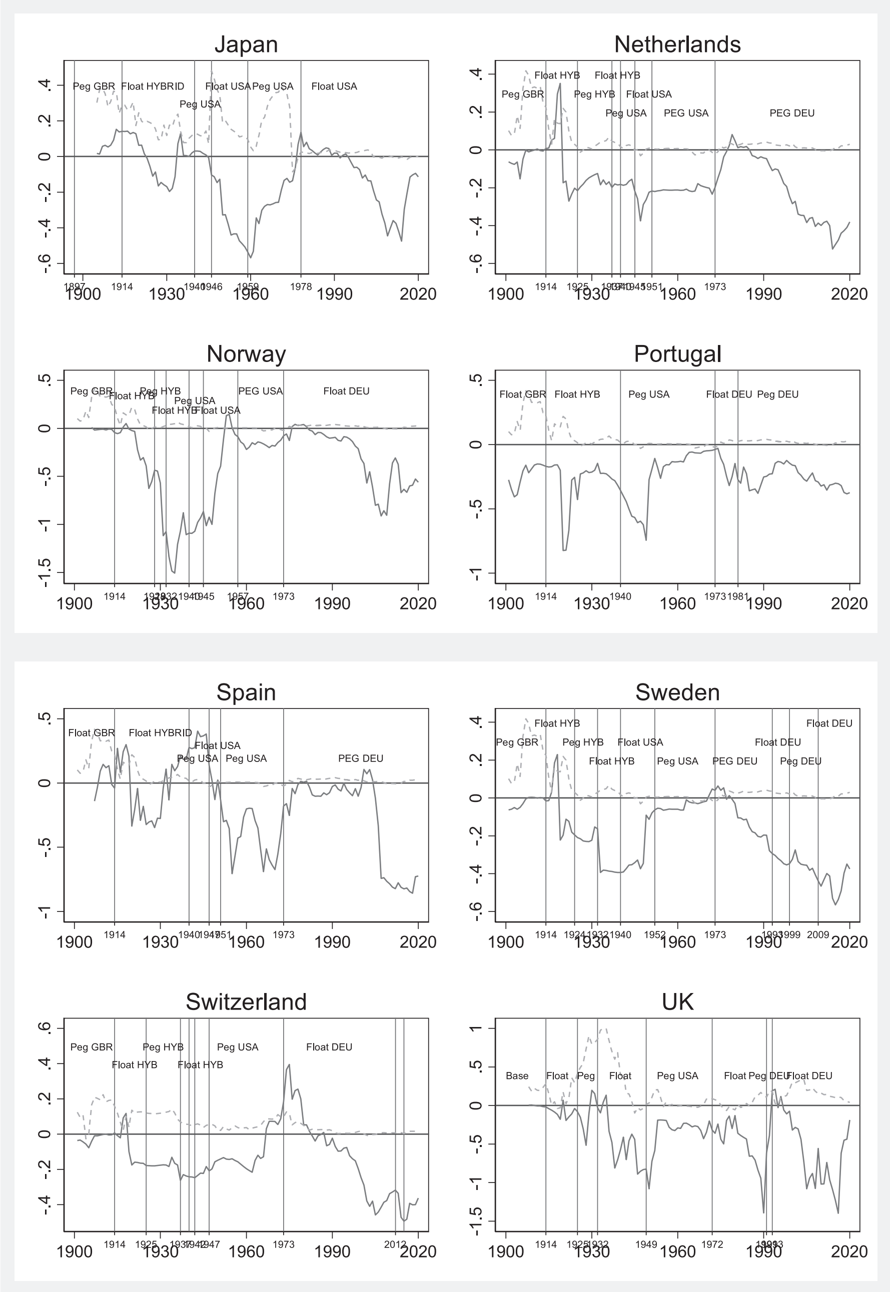

Fig. 3. (*continued*).

exchange rate to the U.S. dollar, for instance, implies zero change in the exchange rate with the U.S. dollar. If PPP holds, the inflation differential must be zero between the two countries. Relative prices must adjust to maintain parity if the exchange rate is pegged and cannot adjust.

We estimate a thirty-year rolling window of adjustments to PPP in the SPMG model of equation (5) to see how adjustments change over time.[10] There is tension in the characterisation of exchange rate regimes (Ilzetzki et al., 2022) so we use the rolling estimator. Using a rolling 30-year window across about 150 years (1871–2020) of the estimation sample, we collect between 113 and 121 rolling estimates (depending on lag length) from each adjustment equation. Figs. 3 and A3 plot the adjustments of exchange rates and relative prices to the parity over time and their *t*-statistics.

[10] We extend the window size to thirty-five years for robustness, as documented in Table A6.

The results are as follows. First, adjusting via exchange rates to PPP strengthens with floating exchange rates, especially since the end of the Bretton Woods (see Fig. 3). Germany and Italy, whose exchange rates were not adjusting to PPP based on the full-sample estimates, adjusted to PPP during since the end of the Bretton Woods. Italy's exchange rate was not adjusting to PPP when it was pegged to the dollar. For countries where we can't be sure about the exchange rate adjustment (Denmark, Japan, Norway, Spain, Switzerland), we also see a strengthening of the exchange rate adjustment towards the end of the sample. This is consistent with Choi and Song (2022), who find that convergence speed rises once exchange rate regimes are accounted for. Eichenbaum et al. (2021) also find that exchange rates adjust to PPP during the floating exchange rate regime.

Second, relative prices adjust to parity when exchange rates do not adjust and vice versa. A regression of exchange rate adjustments on relative price adjustments to parity, country by country, finds the pattern in 14/16 countries (see Table A6). A positive association indicates substitution between both channels of adjustments, which are explained by the opposite signs of the exchange rates (negative) and relative prices adjusting to the parity (positive).

5. Conclusion

Relative purchasing power parity is a long-run theory of exchange rates that says that trade in goods and services would equalise exchange rate changes and inflation differential across countries. Its empirical validity continues to be questioned. We use the System Pooled Mean Group (SPMG) model to study the adjustment of exchange rates and relative prices to PPP in 16 advanced economies to the U.S. dollar and prices over 151 years (1870–2020). The SPMG model and bootstrapped confidence interval accommodate episodic cointegration and cross-country dependence. We detect consensus for long-run relative PPP because we pool the long-run, allow two-way long-run causality and use long-span data.

Declaration of competing interest

The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper.

Data availability

The data is available publicly via the Jordà-Schularick-Taylor Macrohistory Database: https://www.macrohistory.net/database/

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Appendix A. Supplementary data

Supplementary data to this article can be found online at https://doi.org/10.1016/j.jimonfin.2024.103204.

References

Andrews, D.W.K., Chen, H.-Y., 1994. Approximately median-unbiased estimation of autoregressive models. J. Bus. Econ. Stat. 12, 187. https://doi.org/10.2307/1391483.

Auer, R., Burstein, A., Lein, S.M., 2021. Exchange rates and prices: evidence from the 2015 Swiss Franc Appreciation. Am. Econ. Rev. 111, 652–686. https://doi.org/10.1257/AER.20181415.

Balassa, B., 1964. The purchasing-power parity doctrine: a reappraisal. J. Polit. Econ. 72, 584–596. https://doi.org/10.1086/258965.

Boyd, D., Smith, R., 1999. Testing for purchasing power parity: econometric issues and an application to developing countries. Manch. Sch. 67, 287–303. https://doi.org/10.1111/1467-9957.00148.

Burstein, A., Gopinath, G., 2014. International Prices and Exchange Rates, in: Handbook of International Economics, 4th Ed. Elsevier, pp. 391–451. Doi: 10.1016/B978-0-444-54314-1.00007-0.

Choi, J.H., Song, S., 2022. Revisiting the PPP puzzle: Nominal exchange rate rigidity and region of inaction. J. Int. Finan. Markets. Inst. Money 78, 101540. https://doi.org/10.1016/j.intfin.2022.101540.

Chudik, A., Pesaran, M.H., Smith, R.P., 2023. Revisiting the great ratios hypothesis. Oxf. Bull. Econ. Stat 85, 1023–1047. https://doi.org/10.1111/obes.12571.

Eichenbaum, M.S., Johannsen, B.K., Rebelo, S.T., 2021. Monetary policy and the predictability of nominal exchange rates. Rev. Econ. Stud 88, 192–228. https://doi.org/10.1093/restud/rdaa024.

Engel, C., Morley, J., 2001. The Adjustment of Prices and the Adjustment of the Exchange Rate. Cambridge, MA. Doi: 10.3386/w8550.

Friedman, M., 1953. The case for flexible exchange rates. In: Friedman, M. (Ed.), Essays in Positive Economics. The University of Chicago Press, Chicago, pp. 157–203.

Froot, K.A., Rogoff, K., 1995. Chapter 32 Perspectives on PPP and long-run real exchange rates. pp. 1647–1688. Doi: 10.1016/S1573-4404(05)80012-7.

Hansen, B.E., 1999. The Grid Bootstrap and the Autoregressive Model. Rev. Econ. Stat. 81, 594–607. https://doi.org/10.1162/003465399558463.

Harding, T., Stefanski, R., Toews, G., 2020. Boom Goes the Price: Giant Resource Discoveries and Real Exchange Rate Appreciation. Econ. J. 130, 1715–1728. https://doi.org/10.1093/ej/ueaa016.
Ilzetzki, E., Reinhart, C.M., Rogoff, K.S., 2022. Rethinking exchange rate regimes. pp. 91–145. Doi: 10.1016/bs.hesint.2022.02.010.
Imbs, J., Mumtaz, H., Ravn, M.O., Rey, H., 2005. PPP Strikes Back: Aggregation and the Real Exchange Rate. Q. J. Econ 120, 1–43.
Itskhoki, O., Mukhin, D., 2021. Exchange Rate Disconnect in General Equilibrium. J. Polit. Econ. 129, 2183–2232. https://doi.org/10.1086/714447.
Johansen, S., 1991. Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models. Econometrica 59, 1551. https://doi.org/10.2307/2938278.
Jordà, Ò., Schularick, M., Taylor, A.M., 2017. Macrofinancial History and the New Business Cycle Facts. NBER. Macroecon. Annu 31, 213–263. https://doi.org/10.1086/690241.
Lopez, C., Murray, C.J., Papell, D.H., 2013. Median-unbiased estimation in DF-GLS regressions and the PPP puzzle. Appl. Econ 45, 455–464. https://doi.org/10.1080/00036846.2011.605761.
Müller, U.K., Watson, M.W., 2018. Long-Run Covariability. Econometrica 86, 775–804. https://doi.org/10.3982/ecta15047.
Murray, C.J., Papell, D.H., 2002. The purchasing power parity persistence paradigm. J. Int. Econ 56, 1–19. https://doi.org/10.1016/S0022-1996(01)00107-6.
Papell, D.H., Prodan, R., 2020. Long-run purchasing power parity redux. J. Int. Money. Finance 109, 102260. https://doi.org/10.1016/j.jimonfin.2020.102260.
Pesaran, M.H., Smith, R.P., 1998. Structural Analysis of Cointegrating VARs. J. Econ. Surv 12, 471–505. https://doi.org/10.1111/1467-6419.00065.
Rogoff, K., 1996. The purchasing power parity puzzle. J. Econ. Lit 34, 647–668.
Rossi, B., 2005. Confidence intervals for half-life deviations from purchasing power parity. J. Bus. Econ. Stat. 23, 432–442. https://doi.org/10.1198/073500105000000027.
Samuelson, P.A., 1964. Theoretical notes on trade problems. Rev. Econ. Stat 46, 145. https://doi.org/10.2307/1928178.
Steinsson, J., 2008. The dynamic behavior of the real exchange rate in sticky price models. Am. Econ. Rev. 98, 519–533. https://doi.org/10.1257/aer.98.1.519.
Taylor, A.M., Taylor, M.P., 2004. The purchasing power parity debate. J. Econ. Perspect. 18, 135–158. https://doi.org/10.1257/0895330042632744.
Vo, H.L., Vo, D.H., 2022. The purchasing power parity and exchange-rate economics half a century on. J. Econ. Surv 1–34. https://doi.org/10.1111/joes.12504.