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Purchasing power parity (PPP) in developed countries: updated evidence from multivariate ARDL unit root test

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Abstract

This study re-examines the purchasing power parity (PPP) hypothesis for six developed countries, addressing the inconsistent evidence in the existing empirical literature. By applying a newly developed multiple autoregressive distributed lag (MARDL) unit root test, this study updates that both real and real effective exchange rates of the UK, Canada, Australia, Japan, Switzerland, and New Zealand (with the US as a covariate) are non-stationary, I(1) for 2000m1-2023m12, suggesting that the PPP does not hold.

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1. Introduction

The purchasing power parity (PPP) hypothesis is initially articulated by David Ricardo in the early 19th century (see, Ricardo, 1821). It refers to the exchange rate between currencies, which should ensure that the cost of a standard basket of goods and services is the same between countries, thereby equalizing their purchasing power. However, Gustav Cassel formally introduced and popularized it in the early 20th century (Boundi-Chraki & Tomé, 2022). Since then, the PPP hypothesis has been extensively tested on both country-specific and cross-country levels, with different methodologies, but its empirical validity remains inconclusive (for example, Cuddington & Liang, 2000; Kyei-Mensah,2023). Most studies rely on conventional univariate unit root tests (i.e. augmented Dickey-Fuller test) and their nonlinear variants, while the multivariate approach remains unexplored. One major limitation of univariate unit root tests is that it has low power in validating PPP. Using a newly developed multivariate autoregressive distributed lag (MARDL) unit root test (Sam et al., 2024), this study re-examines the PPP hypothesis for both real and real effective exchange rates of six developed countries, namely the UK, Canada, Australia, Japan, Switzerland, and New Zealand, with the US as a covariate. Further confirmation of the Purchasing Power Parity (PPP) hypothesis is crucial for at least two reasons, especially for developed countries. Firstly, PPP offers a more accurate measure of living standards across countries with varying price levels, making comparisons of economic indicators such as GDP and income levels more relevant and precise. Secondly, governments and international organizations refer to PPP data to formulate appropriate policies, allocate resources efficiently, and assess the impact of economic interventions more accurately.

A literature review of selected PPP studies is documented in Section 2. Section 3 details the PPP model, data, and MARDL unit root tests. Sections 4 and 5 elaborates on the results and conclude the study, respectively.

2. Literature Review

Although there have been enormous studies on purchasing power parity (PPP), the existing literature offers inconclusive evidence on PPP. PPP asserts that the real exchange rate reverts to a constant mean which can be tested with a unit root test. If rejecting the null hypothesis which in favor of a level stationary, then there is evidence that PPP holds. Employing Augmented-Dickey–Fuller (ADF) tests on univariate real exchange rates for developed nations typically shows that the null hypothesis of a unit root is seldom rejected (Cuddington and Liang, 2000; Wu and Wu, 2001). A growing body of literature has turned to panel data methods by advocating pooling data across currencies to search for more evidence of PPP. For example, Papell (2002) finds the real exchange rate for a panel of sixteen industrialized countries is stationary, while a unit root is found for a panel of twenty countries.

As pointed out by McNown and Wallace (1989), a cointegrated relation should be formed by the exchange rate and relative price levels if PPP serves as an equilibrium constraint. Hence, cointegration is a *necessary* condition for PPP to meet as a long-run constraint, while non-cointegration implies that PPP does not tend to hold. Employing the Engle-Granger test for cointegration, Taylor and McMahon (1988) obtained strong evidence of long-run PPP. Using

Engle-Granger and Johansen cointegration tests, Taylor (1992) also found that the dollar-sterling exchange rate did hold the long-run PPP from early 1921 until the return of sterling to the gold standard in 1925. Kugler and Lenz (1993) found mixed findings for the post-1973 period of the floating exchange rate, they found that PPP seems to hold in the long-run for six European countries, however, PPP is rejected for the US and the Canadian dollar, the Danish Krone and the Belgian Franc. Using both the Engle-Granger and the Johansen method, Chocholata (2009) does not find support for long-run PPP between Latvia, Slovakia, and the euro area from January 1999 to May 2008. Baharumshah and Ariff (1997) also do not confirm the purchasing power parity (PPP) validity in five Asian countries. Engle (2000) commented that those studies that find evidence in favor of long-run PPP may have drawn incorrect conclusions. His study demonstrates that tests on long-run PPP exhibit significant size biases.

Existing methodology examined the long-run PPP either using the unit root test to examine whether the real exchange rate reverted to the mean or whether the exchange rate and relative price levels should form a cointegrated system using the cointegration test. The MARDL unit root test which was newly developed by Sam et al. (2024) allows to testing of stationarity and cointegration elements in one test. This study fills in this gap by adopting the MARDL unit root test for six major industrial countries using the latest data.

3. Model and Methodology

The PPP Model

Purchasing Power Parity (PPP) hypothesis advances that exchange rates between currencies should change so that identical goods and services cost the same in different countries in the long-run if they priced in a common currency. That is, the exchange rate between two currencies is in equilibrium when their purchasing power is equal in both countries. PPP is initiated by the "*law of one price*" that identical goods and services should trade for the same price in different countries for a single currency, in the absence of transportation costs and trade barriers. PPP hypothesis is commonly delivered in two forms, viz. absolute PPP and relative PPP. The former assumes an equilibrium exchange rate equalizes to relative prices i.e. $E = P/P^*$ where E is the exchange rate, P is the domestic price, and P^* is foreign price. The latter illustrates that any change in the exchange rate between two currencies is their relative inflation rate i.e. $\Delta E/E = \Delta P/P - \Delta P^*/P$. An exchange rate then can be expressed by an equalization that, $EP = EP^*$. The validity of PPP is technically informed by the stationarity of the real exchange rate (*RER*) (Rogoff, 2006), in which $RER = N \times P/P^*$, where N is the nominal exchange rate. If *RER* is one, PPP holds because the purchasing power of the two country's currencies is equivalent when price levels are being considered. Any deviations from one invalidate PPP, suggesting either overvalued or undervalued. Similar implication to the real effective exchange rate (*REER*).

Multivariate autoregressive distributed lag (MARDL) unit root test

Building on the ARDL framework of McNown et al. (2018), Sam et al. (2024) developed a new approach namely the MARDL unit root test in testing a series' stationarity. Conventional unit root tests are known to have low power properties with a higher risk of falling to Type II error. Hansen (1995) explains that a univariate framework often ignores relevant information in explaining multivariate macroeconomic data sets. Adding relevant variables to the model helps to decrease the variance in the model's error, and therefore the OLS estimates would be more precise and statistically powerful with smaller confidence intervals. Hansen shows that his multivariate unit root test, covariate Augmented Dickey-Fuller (CADF), is successful by improving the power of the test dramatically. Nevertheless, Hansen's CADF framework doesn't cover the possibility of cointegrating relationships in a series. Inspired by McNown et al.'s work, the MARDL based on the ARDL framework further improves Hansen's approach by extending its CADF framework to incorporate cointegrating relationships in testing time series stationarity. Through findings, Sam et al. (2024) show that the CADF approach would face a size distortion problem if the process of the time series is cointegrated with the relevant process. Similar to CADF test statistic distribution, the involving test statistics in MARDL, the *t*-test for lagged dependent variable and the *F*-test for lagged independent variables, are complicated. They are a convex combination of the standard normal distribution and the Dickey-Fuller distribution with nuisance parameters dependent. To overcome the uncertainty in determining the distributions, the bootstrap method is used.

To begin with the MARDL unit root test, consider a single ARDL equation:

$$\Delta y_t = c + \beta_1 y_{t-1} + \beta_2 x_{t-1} + \textstyle\sum_{i=1}^{p} \phi_i \Delta y_{t-1} + \sum_{j=1}^{q} \omega_j \Delta x_{t-j} + \epsilon_t \tag{1}$$

where y_t represents the dependent variable, x_t is the covariate variable, Δ indicates the first difference operator, c is a constant term, and ε_t is the error term. The stationarity of y_t is tested as follows: *t*-test of $H_0\colon \beta_1 = 0$ against $H_1\colon \beta_1 < 0$; and *F*-test of $H_0\colon \boldsymbol{\beta}_2 = 0$ against $H_1\colon \boldsymbol{\beta}_2 \neq 0$.

Four possible cases arise during the test:

Case I: If $\boldsymbol{\beta_1 = 0; \beta_2 = 0, y_t \sim I(1)}$.

In Case I, equation (1) reduces to equation (2) and it is a nonstationary process with no cointegration. Equation (2) is similar to the CADF equation not rejecting its null hypothesis of coefficient of y_{t-1}, indicating that y_t is an $I(1)$ process.

$$\Delta y_t = c + \sum_{i=1}^{p} \phi_i \Delta y_{t-1} + \sum_{j=1}^{q} \omega_j \Delta x_{t-j} + \epsilon_t \tag{2}$$

Case II: If $\boldsymbol{\beta_1 < 0; \beta_2 = 0, y_t \sim I(0)}$.

Equation (1) becomes:

$$\Delta y_t = c + \beta_1 y_{t-1} + \sum_{i=1}^{p} \phi_i \Delta y_{t-1} + \sum_{j=1}^{q} \omega_j \Delta x_{t-j} + \epsilon_t \tag{3}$$

in which is similar to CADF equation in rejecting its null hypothesis of coefficient of y_{t-1}, indicating y_t is an $I(0)$ process. The stationarity of y_t is not affected by the presence of the stationary variables Δx_t.

Case III: If $\boldsymbol{\beta_1 = 0; \beta_2 \neq 0, y_t \sim I(2)}$.

Equation (1) will be reduced to:

$$\Delta y_t = c + \beta_2 x_{t-1} + \sum_{i=1}^{p} \phi_i \Delta y_{t-1} + \sum_{j=1}^{q} \omega_j \Delta x_{t-j} + \epsilon_t \tag{4}$$

Case III tells that y_t is a nonstationary $I(2)$ process. The sources of nonstationary in y_t come from the marginal process of itself (y_t), and x_t.

Case IV: If $\boldsymbol{\beta_1 < 0; \beta_2 \neq 0, y_t \sim I(1)}$.
For Case IV the equation has the identical representative as in equation (1):

$$\Delta y_t = c + \beta_1 y_{t-1} + \beta_2 x_{t-1} + \sum_{i=1}^{p} \phi_i \Delta y_{t-1} + \sum_{j=1}^{q} \omega_j \Delta x_{t-j} + \epsilon_t \quad (1)$$

It informs that y_t is nonstationary, and cointegration exists between y_t and x_t. Since x_t is *I*(1), then y_t must also be *I*(1), because both y_t and x_t must have the same order of integration, *I*(d), to be cointegrated.

In short, the testing series is stationary if and only if the *t*-test is significant and *F*-test is insignificant, as in Case II. Other combination results tell that the series is nonstationary. The null distributions of the test statistics are complicated and nuisance parameters dependent. Sam et al. (2024) use the bootstrap method to generate empirical critical values for the *t*- and *F*-tests. Below are the steps to carry out bootstrap to get the *t*- and *F*-tests critical values.

Step 1: Get an optimal ARDL regression that best describes the relationship between the domestic real interest rates, *r*, and the US ex-post real interest rates, r^*:

$$\Delta y_t = \hat{c}_1 + \hat{c}_2 t + \hat{\beta}_1 y_{t-1} + \hat{\beta}_2 x_{t-1} + \sum_{i=1}^{p-1} \hat{\phi}_i \Delta y_{t-i} + \sum_{j=1}^{q-1} \hat{\theta}_j \Delta x_{t-j} + u_t, \quad (5)$$

and estimate with OLS to obtain the parameter estimates.

Step 2: To bootstrap, restrict the test we wish to bootstrap with its null hypothesis into the optimal ARDL(*p*, *q*) regression (5), then obtain *n* observations of OLS estimated residuals. That is, for bootstrapping *t*-test, get *n* observations of OLS estimated residuals of

$$\tilde{u}_t = \Delta y_t - \hat{c}_1 - \hat{c}_2 t - \beta_0 y_{t-1} - \hat{\beta}_2 x_{t-1} - \sum_{i=1}^{p-1} \hat{\phi}_i \Delta y_{t-i} - \sum_{j=1}^{q-1} \hat{\theta}_j \Delta x_{t-j}; \quad (6)$$

For bootstrapping *F*-test, get *n* observations of OLS estimated residuals of

$$\tilde{u}_t = \Delta y_t - \hat{c}_1 - \hat{c}_2 t - \hat{\beta}_1 y_{t-1} - \beta_0 x_{t-1} - \sum_{i=1}^{p-1} \hat{\phi}_i \Delta y_{t-i} - \sum_{j=1}^{q-1} \hat{\theta}_j \Delta x_{t-j}, \quad (7)$$

where $\beta_0 = 0$ is the null of the *t*-test or *F*-test.

Step 3: After obtaining *n* sample size of residuals $\tilde{u}_t$, recenter and rescale these residuals by

$$\ddot{u}_t = \tilde{u}_t - (n - q - 1)^{-1} \sum_t \tilde{u}_t, \quad (8)$$

where *q* is the (maximum) lag length. Save the recentered and rescaled residuals.

Step 4: Resample the recentered and rescaled residuals $\ddot{u}_t$ with replacement to obtain *n* observations of bootstrap $\ddot{u}_t^*$.

Step 5: Generate first-differenced bootstrap observation using the OLS estimates in equation (5). If it is a *t*-test, then

$$\Delta y_t^* = \hat{c}_1 + \hat{c}_2 t + \beta_0 y_{t-1}^* + \hat{\beta}_2 x_{t-1} + \sum_{i=1}^{p-1} \hat{\phi}_i \Delta y_{t-i}^* + \sum_{j=1}^{q-1} \hat{\theta}_j \Delta x_{t-j} + u_t, \quad (9)$$

else if *F*-test, then

$$\Delta y_t^* = \hat{c}_1 + \hat{c}_2 t + \hat{\beta}_1 y_{t-1}^* + \beta_0 x_{t-1} + \sum_{i=1}^{p-1} \hat{\phi}_i \Delta y_{t-i}^* + \sum_{j=1}^{q-1} \hat{\theta}_j \Delta x_{t-j} + u_t. \quad (10)$$

Step 6: Generate the bootstrap observation in level, y_t^* by

$$y_t^* = y_{t-1}^* + \Delta y_t^*. \quad (11)$$

Step 7: Repeat Steps 5 and 6 for *n* times starting from the first to the *n*-th observation to get *n* sample size of y_t^*.

Step 8: Estimate an ARDL regression with similar specification as in equation (5) using the bootstrap observations y_t^*:

$$\Delta y_t^* = \hat{c}_1^* + \hat{c}_2^* t + \hat{\beta}_1^* y_{t-1} + \hat{\beta}_2^* x_{t-1} + \sum_{i=1}^{p-1} \hat{\phi}_i^* \Delta y_{t-i} + \sum_{j=1}^{q-1} \hat{\theta}_j^* \Delta x_{t-j} + \varepsilon_t. \quad (12)$$

Obtain the OLS estimate of $\hat{\beta}_1^*$, if it is *t*-test, or $\hat{\beta}_2^*$, if it is *F*-test, including its corresponding standard errors.

Step 9: Compute the bootstrap test statistic. For the *t*-test, the bootstrap *t*-statistic is computed by

$$t^* = \frac{\hat{\beta}_1^* - \beta_0}{SE(\hat{\beta}_1^*)}; \tag{11}$$

For *F*-test, the bootstrap *F*-statistic is computed by $F^* = \frac{\hat{\beta}_2^* - \beta_0}{SE(\hat{\beta}_2^*)}$, (12)

where SE stands for standard error.

Step 10: Repeat Steps 4 to Step 9 for B times to obtain t_b^* or F_b^*, where $b = 1, \cdots, B$. The ordered bootstrap test statistics are used to construct an empirical bootstrap distribution and determine their critical values from the distribution. The critical value for the *t*-test at α-quantile is given by

$$c_\alpha^* = \max\left\{c : \sum_{b=1}^{B} I\left(t_b^* < c\right) \leq \alpha\right\}, \tag{13}$$

and critical value for *F*-test at $100\%(1-\alpha)$-quantile by

$$c_{1-\alpha}^* = min\{c: \sum_{b=1}^{B} I(t_b^* > c) \leq \alpha\}, \tag{14}$$

where I() is the indicator function, which is 1 given the statement is true and 0 otherwise.

The program code for the multivariate ARDL unit is available at the following link, https://opendata.usm.my/handle/123456789/74723

Data

The real exchange rates (*RER*) and real effective exchange rates (*REER*) of the UK, Canada, Australia, Japan, Switzerland, and New Zealand[1] are monthly data between 2000m1 to 2023m12. The *REER* is obtained from the International Financial Statistics (IFS), the International Monetary Fund (IMF). The *RER* data are constructed manually from the nominal exchange rate (*NER, for example, Yen in US dollar*) and the ratio of prices in the two countries, by $RER = NER \times P/P^*$, where P is the average price of goods in a domestic country, P^* is the average price of good in the US.

The US dollar is being assumed as a covariate variable given its influence on a wide range of global economic outcomes. Figure 1 illustrates both US exchange rates are moving together over the sample period. Hence, *REER* is being considered for a robustness check which is more comprehensive by considering multiple trading partners and their respective trade weights.

[1] The other three G7 member countries, i.e. France, Germany, and Italy are not included in the sample size because these are European Union (EU) member countries which adopts the fixed exchange rates regime.

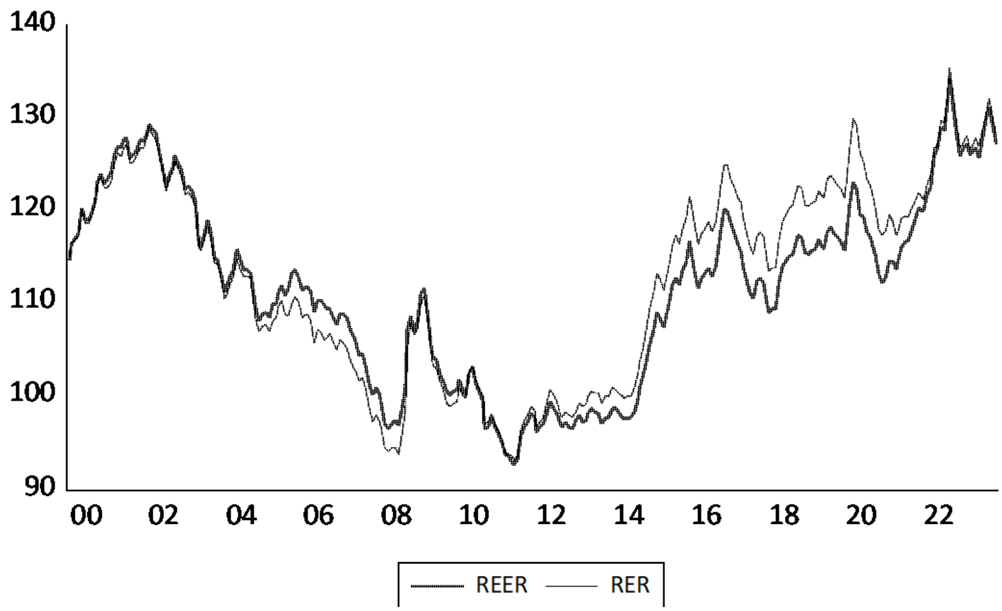

Figure 1. US exchange rates (REER and RER) 2000m1-2023m12

4. The Results

Table 1 reveals that both US exchange rates *RER* and *REER* are non-stationary, $I(1)$ as informed by a set of unit root tests, namely Augmented Dickey-Fuller (ADF), GLS Dickey-Fuller (DF-GLS), Phillips-Perron (PP), and Kwiatkowski, et al. (KPSS) tests. For *RER* and *REER* variables at levels, the respective test statistics fail to reject the null hypothesis of a unit root, while the KPSS test rejects the null hypothesis of stationary. For the first-differenced (Δ) variables, the null hypothesis can be rejected at 1% level, except for KPSS i.e. the null is that the time series variable is stationary, suggesting that the US *RER* and *REER are I(1).*

Table 1. Unit root tests for US exchange rates

	ADF	DF-GLS	PP	KPSS	I(d)
RER	-3.453	-2.573	-3.453	0.739^{***}	
Δ*RER*	-10.939^{***}	-3.703^{***}	-11.005^{***}	0.190	$I(1)$
REER	-3.453	-2.573	-3.453	0.739^{***}	
Δ*REER*	-11.252^{***}	-3.320^{***}	-11.053^{***}	0.216	$I(1)$

Note: *** indicates significance at 1% level.

Table 2 presents the results of the MARDL unit root tests for both *RER and REER* of the six countries - UK, Canada, Australia, Japan, Switzerland, and New Zealand with the US as a covariate. For the *RER* (top panel), the *t*-statistic for all six countries fails to reject the null hypothesis of $\beta_1 = 0$, while the *F*-statistic also fails to reject the null hypothesis of $\beta_2 = 0$. Therefore, both *RER* and *REER* have a unit root (i.e. non-stationary) indicating the absence of PPP for these developed countries. Similar findings are observed for *REER* (bottom panel). These findings are in line with Case I: If $\beta_1 = 0$; $\beta_2 = 0$, $y_t \sim I(1)$ implying that there is a non-stationary process with no cointegration. This is a supportive finding of Cuddington and Liang (2000), Holmes (2001), and Hyrina and Serletis (2010), for example.

Table 2. MARDL unit root tests

Countries	lag	t-statistic	10% critical value	H_0: $\beta_1 = 0$	F-statistic	10% critical value	H_0: $\beta_2 = 0$
Real exchange rate, *RER*							
UK	(0,3)	-1.860	-2.617	Do not reject H_0	0.551	3.863	Do not reject H_0
Canada	(0,5)	-2.542	-2.869	Do not reject H_0	2.741	4.718	Do not reject H_0
Australia	(1,3)	-2.137	-2.894	Do not reject H_0	1.189	4.521	Do not reject H_0
Japan	(2,0)	-2.490	-2.809	Do not reject H_0	2.257	4.156	Do not reject H_0
Switzerland	(0,0)	-2.577	-2.849	Do not reject H_0	0.150	3.942	Do not reject H_0
New Zealand	(0,3)	-2.582	-2.803	Do not reject H_0	0.624	4.038	Do not reject H_0
Real effective exchange rate, *REER*							
UK	(5,0)	-1.293	-2.881	Do not reject H_0	0.016	4.972	Do not reject H_0
Canada	(5,0)	-2.382	-2.859	Do not reject H_0	3.815	4.815	Do not reject H_0
Australia	(5,1)	-0.879	-2.750	Do not reject H_0	0.006	4.675	Do not reject H_0
Japan	(5,0)	-1.476	-2.808	Do not reject H_0	1.386	4.658	Do not reject H_0
Switzerland	(5,0)	-1.361	-2.796	Do not reject H_0	1.458	4.807	Do not reject H_0
New Zealand	(5,0)	-1.105	-2.729	Do not reject H_0	0.075	4.045	Do not reject H_0

Notes: The US data serves as a covariate. The lags are determined by the modified Akaike Information Criterion (AIC). The critical values are computed by using the bootstrap approach with 5,000 replications.

5. Conclusion

This study updates new evidence that the PPP does not hold for six developed countries namely, the UK, Canada, Australia, Japan, Switzerland, and New Zealand for the period 2000m1-2023m12. It is based on both their real and real effective exchange rates are non-stationary, $I(1)$ from the multivariate autoregressive distributed lag (MARDL) unit root test. A few implications include the potential for overvalued or undervalued currencies, leading to exchange rate volatility and misalignment. This can deteriorate countries' global competitiveness, create arbitrage opportunities, and cause imbalances in trade and capital flows. Additionally, it increases the risk and uncertainty of international investments, which affects the formulation of appropriate monetary and fiscal policies, including exchange rate policies.

References

Baharumshah, A. Z. and Ariff Mohamed (1997) "Purchasing power arity in South East Asian countries economies: a cointegration approach" *Asian Economic Journal* **11**(2), 141-153.

Boundi-Chraki, F. and Tomé, J. P. M. (2022) "The purchasing power parity hypothesis tested once again: new empirical evidence for 28 OECD countries" *Investigación Económica* **81**(322), 3–26.

Chocholata, M. (2009) "Purchasing power parity and cointegration: evidence from Latvia and Slovakia" *Ekonomický časopis* **57**(4), 344 – 358.

Cuddington, J. T. and Liang, H. (2000) "Purchasing power parity over two centuries?" *Journal of International Money and Finance* **19**(5), 753-757.

Engle, C. (2000) "Long-run PPP may not hold after all" *Journal of International Economics* **57**, 243-273.

Holmes, M. J. (2001) "New evidence on real exchange rate stationarity and purchasing power parity in less developed countries" *Journal of Macroeconomics* **23**(4), 601-614.

Hyrina, Y. and Serletis, A. (2010) "Purchasing power parity over a century" *Journal of Economic Studies*, **37**(1), 117-144.

Kugler, P., and Lenz, C. (1993) "Multivariate cointegration analysis and the long-run validity of PPP" *Review of Economics and Statistics* **75**(1), 180-184.

Kyei-Mensah, J. (2023) "The long-run validity of PPP in some major advanced and emerging countries using alternative models" *Cogent Economic and Finance* **11**(1), 1-26.

McNown, R., Sam C. Y., and Goh, S. K. (2018) "Bootstrapping the autoregressive distributed lag test for cointegration" *Applied Economics* **50**(13), 1509-1521.

Papell, D. H. (2002) "The great appreciation, the great depreciation, and the purchasing power parity hypothesis" *Journal of International Economics* **57**(1), 51-82.

Ricardo, D. (1821) *On the Principles of Political Economy and Taxation*, Liberty Fund Inc.: Indianapolis.

Rogoff, K. (1996) "The purchasing power parity puzzle" *Journal of Economic Literature* **34**(2), 647-668.

Sam, C. Y., Mcnown R., Goh, S. K., and Goh, K. L. (2024) "A multivariate autoregressive distributed lag unit root test" *Studies in Economics and Econometrics* **49**(1), 17-33.

Taylor, M.P. (1992) "Dollar-Sterling exchange rate in the 1920s: purchasing power parity and the Norman conquest of $4.86" *Applied Economics* **24**(8), 803-11.

Taylor, M. P. and McMahon, P. C. (1988) "Long-run purchasing power parity in the 1920s" *European Economic Review* **32**, 179-97.

Wu, J. L., and Wu, S. (2001) "Is purchasing power parity overvalued?" *Journal of Money, Credit and Banking* **33**(3), 804-812.