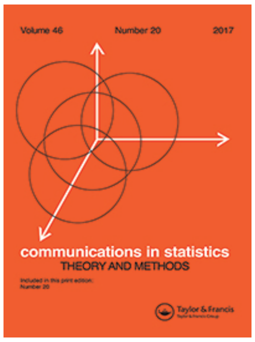


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# Revisiting purchasing power parity in BRICS countries using more powerful quantile unit-root tests with stationary covariates

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**ABSTRACT**
This study contributes to this line of research by determining whether purchasing power parity (PPP) holds true in BRICS countries (i.e., Brazil, Russia, India, China, and South Africa). We test the hypothesis of PPP in real exchange rate using a more powerful quantile unit-root test with stationary covariates. Our empirical findings indicate a support of PPP for all BRICS countries under study. Our study has important policy implications for the government of the BRICS countries conducting exchange rate policy through PPP.

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BRICS countries; purchasing power parity; quantile unit-root test; stationary covariates.

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## 1. Introduction

In this study, we revisit purchasing power parity (hereafter, PPP) for BRICS countries over 1995M1 to 2015M2, using a newly developed quantile unit-root test with stationary covariates as proposed by Galvao (2009). Investigating whether PPP holds true is critical not only for empirical researchers but also for policymakers. PPP states that due to arbitrage activities in the international commodities market, the real exchange rates that combine the movements of relative prices with nominal exchange rates are expected to return to a constant equilibrium value in the long run (Bahmani-Oskooee et al., 2014). This means that real exchange rate will be a mean-reverting process—stationary in the long run. In particular, a non stationary real exchange rate indi that there is no long-run relationship between nominal exchange rate and domestic and foreign prices, thereby invalidating the PPP hypothesis (Lin et al., 2011). As such, PPP cannot be used to determine the equilibrium exchange rate, and an invalid PPP also disqualifies the monetary approach to exchange rate determination, which requires PPP to hold true. Empirical evidence on the stationarity of real exchange rates is abundant but inconclusive thus far. Previous studies by MacDonald and Taylor (1992), Rogoff (1996), Taylor and Sarno (1998), Lothian and Taylor (2000, 2008), Sarno and Taylor (2002), Taylor and Taylor (2004), Bahmani-Oskooee and Hegerty (2009), Chang and Tzeng (2013), He and Chang (2013), He et al. (2014), and Bahmani-Oskooee et al. (2014) have provided in-depth information on the theoretical and empirical aspects of PPP and the real exchange rate.

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However, the above tests usually focus on the average behavior of real exchange rate without considering the influence of various sizes of shocks on real exchange rate. In other words, the speed of adjustment in real exchange rate toward its equilibrium is usually assumed to be constant, no matter how big or what sign the shock is. As a result, the commonly used conventional unit-root tests possibly lead to a widespread failure in the rejection of unit-root null hypothesis real exchange rate. This paper intends to deal with the above deficiency by employing a newly developed more powerful quantile unit-root test with stationary covariates in Galvao (2009) to enhance estimation accuracy.[1]

This study contributes to this line of research by determining whether PPP holds true in BRICS countries (i.e., Brazil, Russia, India, China, and South Africa). We test the hypothesis of PPP in real exchange rate using a more powerful quantile unit-root test with stationary covariates, proposed by Galvao (2009). Our empirical findings indicate a support of PPP for all BRICS countries under study. Our study has important policy implications for the government of the BRICS countries conducting exchange rate policy through PPP.

The remainder of this paper is organized as follows. Section 2 presents the data used in our study. Section 3 first briefly describes the quantile unit-root test with stationary covariates proposed by Galvao (2009) and then presents the empirical results and its policy implications. Section 4 concludes the paper.

## 2. Data

This empirical analysis covers BRICS countries: Brazil, Russia, India, China, and South Africa. We employed monthly data in our empirical study and the time span is from January 1995 to February 2015. All consumer price indices, consumer price index (CPI) (based on 2005 = 100), and nominal exchange rates relative to the US dollar data, respectively, are taken from the datastream (see Table 1). Each of the consumer price index and nominal exchange rate series was transformed into natural logarithms before performing the econometric analysis. Testing for the PPP against the USA is based on the argument that internal foreign exchange markets are mostly dollar dominated. In addition, funds for economic reconstructions are being provided by the US sponsored institutions (Bahmani-Oskoee et al., 2015).

## 3. Methodology, empirical results, and policy implications

### 3.1. Quantile unit-root test

Let $\mathrm{re}_t$ denote the log of real exchanger rate in our case and $\varepsilon_t$ a serially uncorrelated error term. The AR($q$) process of real exchange rate at quantile $\tau$ can be written as

$$Q_\tau(\mathrm{re}_t|\mathrm{re}_{t-1},\ldots,\mathrm{re}_{t-q}) = \alpha(\tau)\mathrm{re}_{t-1} + a(\tau) + \sum_{i=1}^{q-1}\phi_i(\tau)\Delta\mathrm{re}_{t-i} \tag{1}$$

The test has been extended by Galvao (2009) to include deterministic components which is essential for unit-root tests of drifting time series like real exchange rate in our case. Following Galvao (2009), we can extend Equation (1) to include stationary covariates for unit-root

[1] Including stationary covariates in the regression equation is one promising approach for improving the power of unit root tests, as proposed by Hansen (1995). Hansen (1995) showed that additional information contained in stationary covariates that are correlated with the series can be exploited to obtain the covariate augmented Dickey-Fuller (CADF) test that has higher power than the ADF test.

**Table 1.** Summary statistics of real exchange rate.

| | Brazil | Russia | India | China | South Africa |
|---|---|---|---|---|---|
| Mean | 0.8246 | 3.7238 | 4.0372 | 2.0038 | 2.1031 |
| Median | 0.7257 | 3.6856 | 4.0935 | 2.0359 | 2.0949 |
| Maximum | 1.6227 | 4.3535 | 4.2433 | 2.1646 | 2.6936 |
| Minimum | 0.4141 | 3.2625 | 3.7337 | 1.7519 | 1.7487 |
| Std. dev. | 0.2857 | 0.3245 | 0.1415 | 0.1241 | 0.1807 |
| Skewness | 0.6845 | 0.4673 | − 0.4931 | − 0.6238 | 0.6011 |
| Kurtosis | 2.3422 | 1.9774 | 1.8821 | 2.0818 | 3.8523 |
| Jarque–Bera | 23.2585 | 19.3529 | 22.4064 | 24.1914 | 21.9008 |
| Probability | 0.000009 | 0.000063 | 0.000014 | 0.000006 | 0.000018 |
| Observations | 242 | 242 | 242 | 242 | 242 |

testing. Galvao (2009) has proofed that his proposed model has more power than standard quantile unit test model. We can express his model as

$$Q_\tau(\mathrm{re}_t|\mathrm{re}_{t-1},\dots,\mathrm{re}_{t-q}) = \alpha(\tau)\mathrm{re}_{t-1} + a(\tau) + \sum_{i=1}^{q-1}\phi_i(\tau)\Delta\mathrm{re}_{t-i}. + \sum_{I=-q1}^{-q2}\gamma_{\mathrm{I}}(\tau)\Delta\mathrm{X}_{t-I} + F_{\mathrm{u}}^{-1}(\tau) \tag{2}$$

where $F_u$ denotes the common distribution function of errors. Let $a(\tau) = a + \mathrm{F}_{\mathrm{u}}^{-1}(\tau)$ define $Z = (1, \Delta\mathrm{re}_{t-1}, \Delta\mathrm{re}_{t-2}, \dots, \Delta\mathrm{re}_{t-q+1}, \Delta\mathrm{X}_{t-\mathrm{q}2}\dots., \Delta\mathrm{X}_{t+\mathrm{q}1})^T$. Here we can define $\Delta X_{t-\mathrm{I}}$ as stationary covariates (such as stock returns and inflation rate in our case). By estimating Equation (2) at different quantiles $\tau \in (0, 1)$, we can get a set of estimates of the persistence measure as $\alpha(\tau)$. We can test $\alpha(\tau) = 1$ at different values of $\tau$ to analyze the persistence of the exchange rate impact of positive and negative shocks and shocks of different magnitude using the quantile autoregression-based unit-root test proposed by Koenker and Xiao (2004).

Let $\widehat{\alpha}(\tau)$ be the quantile regression estimator. To test $H_0 : \alpha(\tau) = 1$, we use the $t$-stat for $\widehat{\alpha}(\tau)$ proposed by Koenker and Xiao (2004) which can be written as

$$t_n(\tau) = \frac{f(\widehat{F^{-1}(\tau)})}{\sqrt{\tau(1-\tau)}}(\mathrm{re}'_{-1}M_Z\mathrm{re}_{-1})^{1/2}(\widehat{\alpha}(\tau) - 1) \tag{3}$$

where $f(u)$ and $F(u)$ are the probability and cumulative density functions of $\varepsilon_t$, $\mathrm{re}_{-1}$ is the vector of lagged log-real exchange rate, and $M_z$ is the projection matrix onto the space orthogonal to $Z = (1, \Delta\mathrm{re}_{t-1}, \Delta\mathrm{re}_{t-2}, \dots, \Delta\mathrm{re}_{t-q+1}, \Delta\mathrm{X}_{t-\mathrm{q}2}\dots., \Delta\mathrm{X}_{t+\mathrm{q}1})^T$. We use the results derived by Koenker and Xiao (2004) and Galvao (2009) to find the critical values of $t_n(\tau)$ for different quantile levels. We can estimate $f(F^{-1}(\tau))$ following the rule given in Koenker and Xiao (2004) and Galvao (2009). Besides allowing for asymmetric effects of shocks on real exchange rate, an important advantage of QAR-based unit-root tests over standard unit-root tests is that they have more power (Koenker and Xiao, 2004; Galvao, 2009).

In contrast, a more complete inference of the unit-root process based on the quantile approach involves exploring the unit-root property across a range of quantiles. To this end, both Koenker and Xiao (2004) and Galvao (2009) suggest the quantile Kolmogorov–Smirnov (QKS) test, which is given as

$$\mathrm{QKS} = \sup_{\tau\in\Gamma}|t_n(\tau)| \tag{4}$$

where $t_n(\tau)$ is given by Equation (3) and $\Gamma = (0.1, 0.2, \dots..0.9)'$ in our latter applications. In other words, we first calculate $t_n(\tau)$ for all $\tau_s$ in $\Gamma$, and then construct the QKS test statistic by selecting the maximum value across $\Gamma$. While the limiting distributions of both $t_n(\tau)$ and

**Table 2.** Univariate unit-root tests.

| | Level | | | 1st difference | | |
|---|---|---|---|---|---|---|
| | ADF | PP | KPSS | ADF | PP | KPSS |
| Brazil | – 1.490(2) | – 1.571 (5) | 0.454 (12)$^{*}$ | – 10.073 (1)$^{***}$ | – 13.587 (0)$^{***}$ | 0.175 (4) |
| Russia | – 1.855(1) | – 1.827 (7) | 1.236 (12)$^{***}$ | – 9.861(4)$^{***}$ | – 10.556 (2)$^{***}$ | 0.158 (6) |
| India | – 0.804 (5) | – 0.752 (0) | 1.504 (12)$^{***}$ | – 12.348 (6)$^{***}$ | – 12.348 (6)$^{***}$ | 0.154 (0) |
| China | – 1.016 (12) | 0.534 (4) | 1.186 (12)8$^{***}$ | – 10.366 (2)$^{***}$ | – 10.366 (2)$^{***}$ | 0.473 (5)$^{***}$ |
| South Africa | – 2.427(1) | – 2.225 (4) | 0.283 (11) | – 11.347 (0)$^{***}$ | – 11.13 (11)$^{***}$ | 0.108 (4) |

$^{***}$, $^{**}$, and $^{*}$ indicate significance at the 0.01, 0.05, and 0.1 levels, respectively. The number in parenthesis indicates the lag order selected based on the recursive $t$-statistic, as suggested by Perron (1989). The number in the brackets indicates the truncation for the Bartlett Kernel, as suggested by the Newey–West test (1987).

QKS tests are non standard, both Koenker and Xiao (2004) and Galvao (2009) also suggest the use of a resampling (number of bootstrap = 10,000 in our case) procedure to approximate their small-sample distributions. Interested readers can refer to Koenker and Xiao (2004) and Galvao (2009) for more detailed description.

### 3.2. Empirical results and policy implications

For comparison purpose, we also incorporate three conventional unit-root tests: augmented Dickey-Fuller (ADF), Phillips and Perron (PP), and Kwiatkowski-Phillips-Schmidt-Shin (KPSS) tests. The results in Table 2 clearly indicate that both the ADF and PP tests fail to reject the null of non stationary real exchange rate for all BRICS countries. KPSS test gets similar results. This result is consistent with those of Lu and Chang (2011), Chen et al. (2011), Chang (2012), Chang et al. (2012), Chang et al. (2012), Su et al. (2012), and Bahmani-Oskooee

**Table 3.** Empirical results of quantile estimation and unit-root tests for each quantile (taking into account stationary covariates—-stock returns)—-Galvao (2009).

| | $\tau$ | 0.1 | 0.2 | 0.3 | 0.4 | 0.5 | 0.6 | 0.7 | 0.8 | 0.9 |
|---|---|---|---|---|---|---|---|---|---|---|
| Brazil | $\alpha_1(\tau)$ | 0.9573$^{***}$ | 0.9809$^{*}$ | 0.9842 | 0.9872$^{*}$ | 0.9922 | 0.9979 | 1.0058 | 1.0108 | 1.0195 |
| | $\hat{\delta}^2$ | 0.2030 | 0.1815 | 0.6414 | 0.5497 | 0.1856 | 0.3166 | 0.0295 | 0.0019 | 0.0289 |
| | Half-lives | 15.884 | 35.943 | n.a. | 53.805 | n.a. | n.a. | n.a. | n.a. | n.a. |
| | QKS for quantiles of 10%–90%: 4.5415$^{***}$(0.003) [1, 1, 1] | | | | | | | | | |
| Russia | $\alpha_1(\tau)$ | 1.0020 | 1.0027 | 0.9999 | 0.9981 | 0.9965 | 0.9885 | 0.9850$^{**}$ | 0.9829 | 0.9760 |
| | $\hat{\delta}^2$ | 0.0981 | 0.0050. | 0.0297. | 0.0141. | 0.1256. | 0.0220 | 0.0041 | 0.1688 | 0.00001. |
| | Half-lives | n.a. | n.a. | n.a. | n.a. | n.a. | n.a. | 45.862. | n.a. | n.a. |
| | QKS for quantiles of 10%–90%: 2.1463 [12, 3, 6] | | | | | | | | | |
| India | $\alpha_1(\tau)$ | 1.0170 | 1.0179 | 1.0146 | 1.0097 | 1.0034 | 0.9931 | 0.9776$^{**}$ | 0.9710$^{**}$ | 0.9741$^{*}$ |
| | $\hat{\delta}^2$ | 0.0325 | 0.1802 | 0.2975 | 0.3915 | 0.2499 | 0.4339 | 0.2083 | 0.0332 | 0.1375 |
| | Half-lives | n.a. | n.a. | n.a. | n.a. | n.a. | n.a. | 30.596 | 23.553 | 26.414 |
| | QKS for quantiles of 10%–90%: 2.3426 [1, 1, 1] | | | | | | | | | |
| China | $\alpha_1(\tau)$ | 0.9850$^{**}$ | 0.9917 | 0.9969 | 0.9987 | 0.9983 | 0.9991 | 1.0007 | 0.9976 | 1.0042 |
| | $\hat{\delta}^2$ | 0.0155 | 0.4764 | 0.3895 | 0.7893 | 0.5220 | 0.4103 | 0.7010 | 0.2097 | 0.0519 |
| | Half-lives | 45.862 | n.a. | n.a. | n.a. | n.a. | n.a. | n.a. | n.a. | n.a. |
| | QKS for quantiles of 10%–90%: 2.1062 [12, 1, 1] | | | | | | | | | |
| South Africa | $\alpha_1(\tau)$ | 0.9247$^{***}$ | 0.9452$^{***}$ | 0.9541$^{***}$ | 0.9662$^{***}$ | 0.9743$^{**}$ | 0.9761$^{**}$ | 0.9808 | 0.9918 | 0.9816 |
| | $\hat{\delta}^2$ | 0.3041 | 0.1780 | 0.0035 | 0.0001 | 0.0172 | 0.0689 | 0.1663 | 0.1765 | 0.4697 |
| | Half-lives | 8.854 | 12.299 | 14.752 | 20.158 | 26.623 | 28.654 | | | |
| | QKS for quantiles of 10%–90%: 4.0012$^{***}$(0.005) [1, 1, 1] | | | | | | | | | |

$^{**}$and $^{***}$ denote significance at 5% and 1% levels, respectively. Numbers in parenthesis denote bootstrap $p$-values with the bootstrap replications set to be 10,000. For $\alpha_1(\tau)$, the unit-root null is examined with the $t_n(\tau)$ statistic. The lag lengths $p$ and $q$ are selected based on robust Schwarz information criterion as suggested by Galvao (2009) with a maximum lag set to be 12.

**Table 4.** Empirical results of quantile estimation and unit-root tests for each quantile (taking into account stationary covariates—inflation rates)—Galvao (2009).

| | $\tau$ | 0.1 | 0.2 | 0.3 | 0.4 | 0.5 | 0.6 | 0.7 | 0.8 | 0.9 |
|---|---|---|---|---|---|---|---|---|---|---|
| Brazil | $\alpha_1(\tau)$ | 0.946*** | 0.952*** | 0.966** | 0.971** | 0.991 | 1.006 | 1.023 | 1.035 | 1.023 |
| | $\hat{\delta}^2$ | 0.113 | 0.116 | 0.227 | 0.643 | 0.0943 | 0.0016 | 0.1435 | 0.2879 | 0.0649 |
| | Half-lives | 12.486 | 14.091 | 20.038 | 23.553 | n.a. | n.a. | n.a. | n.a. | n.a. |
| | QKS for quantiles of 10%–90%: 4.718** (0.046) [2, 1, 1] | | | | | | | | | |
| Russia | $\alpha_1(\tau)$ | 1.0217 | 1.0103 | 1.0042 | 0.9985 | 0.9941 | 0.9887** | 0.9793*** | 0.9701*** | 0.9655** |
| | $\hat{\delta}^2$ | 0.4419 | 0.9125 | 0.4488 | 0.5125 | 0.2154 | 0.0800 | 0.0736 | 0.3746 | 0.0415 |
| | Half-lives | n.a. | n.a. | n.a. | n.a. | n.a. | 60.993 | 33.137 | 22.834 | 19.742 |
| | QKS for quantiles of 10%–90%: 3.0336** (0.039) [5, 1, 3] | | | | | | | | | |
| India | $\alpha_1(\tau)$ | 1.038 | 1.0279 | 1.0125 | 1.0001 | 0.9915 | 0.9808** | 0.9784** | 0.9807 | 0.9465*** |
| | $\hat{\delta}^2$ | 0.5857 | 0.7669 | 0.8561 | 0.2419 | 0.1598 | 0.1789 | 0.2658 | 0.0033 | 0.0782 |
| | Half-lives | n.a. | n.a. | n.a. | n.a. | n.a. | 35.753 | 31.742 | n.a. | 12.606 |
| | QKS for quantiles of 10%–90%: 2.8847* (0.095) [2, 1, 1] | | | | | | | | | |
| China | $\alpha_1(\tau)$ | 1.0069 | 1.0065 | 1.0066 | 1.0055 | 1.0052 | 1.0048 | 1.0029 | 1.0011 | 0.9951* |
| | $\hat{\delta}^2$ | 0.4235 | 0.9574 | 0.3186 | 0.1973 | 0.1908 | 0.0843 | 0.2506 | 0.4127 | 0.1368 |
| | Half-lives | n.a. | n.a. | n.a. | n.a. | n.a. | n.a. | n.a. | n.a. | 141.111 |
| | QKS for quantiles of 10%–90%: 3.1253* (0.043) [1, 1, 1] | | | | | | | | | |
| South Africa | $\alpha_1(\tau)$ | 0.9443*** | 0.9518*** | 0.9641** | 0.9748* | 0.9800 | 0.9721** | 0.9871 | 0.9943 | 0.9987 |
| | $\hat{\delta}^2$ | 0.3952 | 0.3275 | 0.2047 | 0.0049 | 0.0008 | 0.0563 | 0.2206 | 0.2021 | 0.4237 |
| | Half-lives | 12.094 | 14.031 | 18.959 | 27.157 | n.a. | 24.495 | n.a. | n.a. | n.a. |
| | QKS for quantiles of 10%–90%: 4.3727** (0.005)[1, 1, 1] | | | | | | | | | |

**and *** denote significance at 5% and 1% levels, respectively. Numbers in parenthesis denote bootstrap $p$-values with the bootstrap replications set to be 10,000. For $\alpha_1(\tau)$, the unit-root null is examined with the $t_n(\tau)$ statistic. The lag lengths $p$ and $q$ are selected based on robust Schwarz information criterion as suggested by Galvao (2009) with a maximum lag set to be 12.

et al. (2014), indicating that PPP does not hold true in most BRICS countries, when conventional unit-root tests are conducted.

Due to the deficiency of conventional unit-root test, in the following we employ a newly developed more powerful quantile unit-root test with stationary covariates proposed by Galvao (2009) to enhance its estimation accuracy. Because the choice of covariates is limited only by the fact that they must be stationary and correlated with the shocks to the real exchange rate. We follow Elliott and Pesavento (2006), Su et al. (2012), and Liu and Chang (2013) to choose the stationary covariates according to economic theory. In our study, we use both inflation and stock returns. Results for these two stationary covariates are reported in Tables 3 and 4, respectively. Based on the results from Tables 3 and 4, we can see that PPP holds true for all BRICS countries, with the exception of Russia, India, and China when stock returns are served as stationary covariates based on QKS. However, we find certain quantile's support of PPP for Russia (at quantile 0.7), India (at quantiles between 0.7 and 0.9), and China (at quantile 0.1). Interesting is that we find strong evidence in support of PPP when inflation rates are served as stationary covariates compared to stock returns. Tables 3 and 4 also calculate half-life of a shock for these BRICS countries. We find that the estimated half-life based on quantile autoregressive model is about 8.54–53.805 months (8 months to 4.5 years) when stock returns are served as stationary covariates and about 12.094–60.993 months (13 months to 5 years) when inflation rates are served as stationary covariates.

One major policy implication of our study is that the validity of using PPP to determine the equilibrium exchange is unambiguous for all BRICS countries. The governments of the BRICS countries (i.e., Brazil, Russia, India, China, and South Africa) can use PPP to predict exchange rate that determines whether a currency is over or undervalued and experiencing difference between domestic and foreign inflation rates. This also means that the central banks

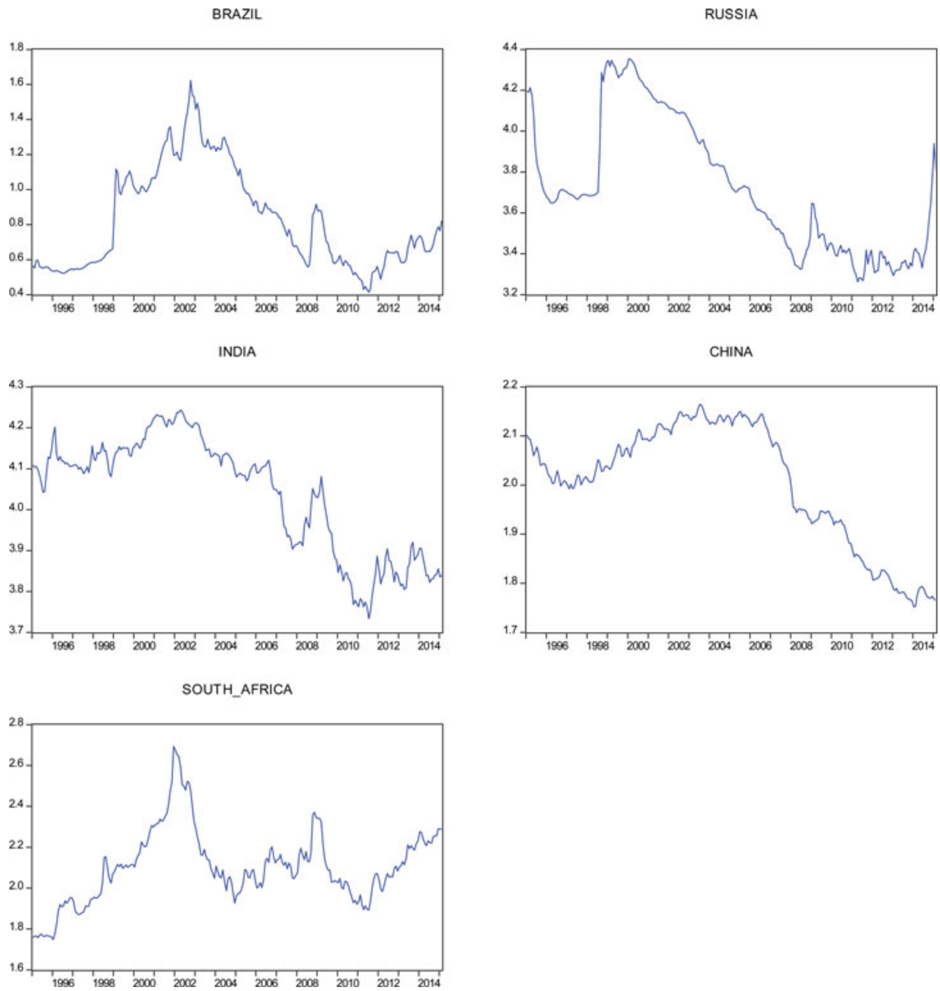


**Figure 1.** Plot of real exchange rates of the BRICS countries.

of the BRICS countries can use PPP to conduct exchange rate policy to determine equilibrium exchange rate for their own countries because real exchange rate is stationary (or PPP holds true). These results imply that the choices and effectiveness of the monetary policies in BRICS economies are highly influenced by external factors that originate from the USA. As can be seen from Figure 1, the central banks of the BRICS countries act accordingly to PPP to appreciate their own currencies against the US dollars to reply the Quantitative Monetary Policy imposed by the US government due to sub-crime crisis during the period of 2007–2009 and the adjustment seems to be non linear and asymmetric. As shown in Table 4, the coefficients are significant at lower quantiles for both Brazil and South Africa, and are significant at higher quantiles for Russia, India, and China, respectively. Nevertheless, reaping unbounded gains from arbitrage in traded goods is not possible in these BRICS countries.

## 4. Conclusions

We revisit the PPP hypothesis for BRICS countries over 1995M1 to 2015M2, using a more powerful quantile unit-root test with stationary covariates as proposed by Galvao (2009). Our empirical findings indicate a support of PPP for all the BRICS countries under study.

Our study has important policy implications for the government of these BRICS countries conducting exchange rate policy through PPP.

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