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Revisiting Purchasing Power Parity for India using threshold cointegration and nonlinear unit root test

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Abstract This study examines the Purchasing Power Parity (PPP) hypothesis in case of India for her five major trading partners over the period of $1991M_1$–$2009M_2$. The study used the DF-GLS unit root test and threshold autoregressive (TAR) model as well as momentum-TAR (M-TAR) models for empirical analysis. However, we relied on TAR and MTAR models based cointegration tests to draw conclusions because of their superiority to traditional cointegration techniques as these models have limit cycles, amplitude dependent frequencies, and jump phenomena. These models are capable of producing asymmetric limit cycles and are suitable for time series data. Our empirical exercise reveals that PPP hypothesis does not exist for all major trading partners in case of India. This reveals that intermediate goods face high barriers to trade in this sampled countries. This supports the argument that Indian government has not been able to strike out the proper balance between flexibility and stability between real bilateral exchange rates and thus unable to maintaining confidence in the domestic currency that has been evident from the recent fall of rupee in relation to the US dollar.

Keywords PPP · Nonlinearity · Unit root · Threshold autoregressive (TAR) model · India

JEL classification F31 · F15

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1 Introduction

Exchange rate is an important instrument to determine international trade and finance following financial liberalizations and globalization epoch in case of emerging economy such as India. Exchange rate movements influence the profitability of multinationals and in resulting, exchange exposure of enterprises and institutions is increased. The enterprises and financial institutions may use stable exchange rate as a tool to assess the performance of investment ventures, financing, hedging as well as reducing their operational risks (Nieh and Wang 2005; Khan and Qayyum 2007). The movements in exchange rate significantly affect the macroeconomic variables such as interest rate, inflation, wages, unemployment, domestic production etc. This may create macroeconomic disequilibrium and local currency may be devalued to correct external imbalances (Parikh and Williams 1998).

The concept of Purchasing Power Parity is highly debated and controversial issue in international finance through which we can assess the long run stable exchange rate. The PPP theory was advanced by Cassel (1916, 1918) reveals that nominal exchange rate and price levels are equal between two economies under free trade conditions. The assumption of PPP theory is that changes in nominal exchange rate counterbalance the movements in relative prices keeping the equilibrium exchange rate constant in long run. The PPP theory is based on the law of one price (LOP) in integrated and competitive markets assuming that world is risk neutral. The PPP theory is considered as an important tool that determines the exchange rate. It is also helpful in evaluating the financial liberalization and structural adjustment policies forced by International Monetary Fund and World Bank. The PPP theory provides guidelines to monetary authorities in choosing money supplying and inflation targeting. Furthermore, PPP is the best criterion to judge whether there is overvaluation or undervaluation in exchange rate with reference to long run equilibrium path of exchange rate. This shows that proper understanding of factors affecting exchange rate is helpful to policy makers to articulate a comprehensive exchange rate policy to sustain balance of trade and hence balance of payment in long run. Despite its importance, PPP theory remains a controversial issue.

There are a number of studies extensively examining the long-run relationship between nominal exchange rate and relative price postulated by the PPP hypothesis by using various types of cointegration and unit root tests. However, most of the studies belong to the developed and developing countries context and there is relatively less work done in the area of economics like India where fast transition is taking place. Therefore, the present study is pioneering effort while determining whether PPP theory holds for India's real exchange rate relative to a sample of her major trading partner countries (Hong Kong, Saudi Arabia, Singapore, UK and US) or not by applying DF-GLS unit root test and threshold autoregressive (TAR) model and momentum-TAR (M-TAR) models. These countries have been selected as they serve an importance destination in Indian exports. Indian exports to the top partners have been presented in the following Table 1.[1]

[1] Table 1 has been sourced from: http://www.infodriveindia.com/export-import/trade-statistics/trading-partners.aspx.

Table 1 Indian Exports to Top Partners

Rank	Countries name	Apr 2005–Feb 2006		Apr 2006–Feb 2007		% Growth (INR in crores)	% Growth (US$ in millions)	% Share (INR in crores)	% Share (US$ in millions)
		Export value of goods		Export value of goods					
		(INR in crores)	(US$ in millions)	(INR in crores)	(US$ in millions)				
1	India export to—USA	68,900.84	15,569.19	76,803.05	16,916.26	11.47	8.65	15.05	15.05
2	India export to—U. Arab Emts.	33,724.33	7,620.53	49,090.65	10,812.47	45.56	41.89	9.62	9.62
3	India export to—China PRP	25,587.57	5,781.90	32,937.39	7,254.63	28.72	25.47	6.45	6.45
4	India export to—Singapore	21,141.93	4,777.34	24,616.62	5,421.94	16.44	13.49	4.82	4.82
5	India export to—UK	20,093.80	4,540.50	22,838.56	5,030.31	13.66	10.79	4.47	4.47
6	India export to—Hong Kong	17,522.76	3,959.53	18,366.82	4,045.39	4.82	2.17	3.60	3.60

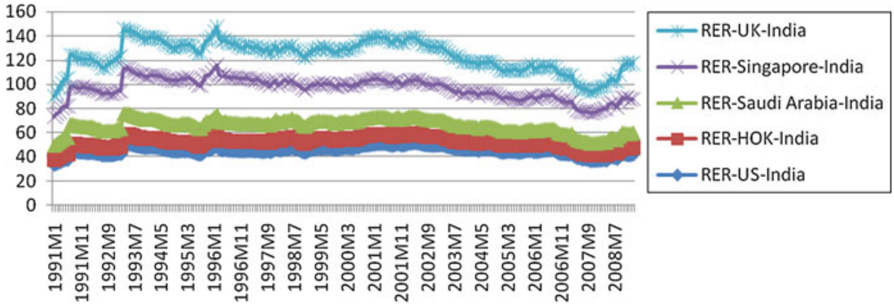

Fig. 1 Time series plots of the India's real exchange rate with major trading partners

Table 1 clearly ranks Indian top export partners wherein USA, United Arab Emirate, China, Singapore, UK and Hong Kong holds at first, second, third, fourth, fifth and sixth place respectively. Since a countries export performance depends upon the exchange rate of Indian rupee to the currency of the respective country. We have plotted the real exchange rate of Indian rupee to the currency of the respective top trading countries in Fig. 1, which depict nonlinear feature. Further, real bilateral exchange rate historical trend plots show high fluctuations in Indian exchange rate with UK and Singapore.[2]

India provides an interesting arena to research particularly because of its remarkable economic progress over the past two decades. Second, as annual report (2009–10) of Government of India, Ministry of Commerce and Industry, department of commerce mentions that Indian economy has been relatively less affected by the global crisis witnessed in the whole world recently. Report further states that in fact, India, along with China, is one of the growth engines in facilitating a faster turnaround of the global economy.[3] Thirdly, India's foreign exchange reserve has been more than the required level from a long period. Fourthly, as mentioned in the "Review of Trade Policies of India's Major Trading Partners (2009)", India's current share of the world merchandise trade is 1.1 % and services trade is 2.7 % and in merchandise trade, India's target is to reach 5 % of the world trade by 2020. Further, the period, which is chosen in this analysis, is post reform period because of few reasons. From 1975 to 1991, the Indian rupee had been adjustably pegged to a currency basket of major trading partners therefore this period of exchange rate was pegged exchange rate system. Further, India's extensive economic liberalization begins in 1991 and India experienced structural changes in the exchange rate regime during the first half of the 1990s with the official devaluation of the rupee in July 1991, before that, the exchange rate was fixed by the central bank, the Reserve Bank of India (RBI). However, if we look into literature we find that studies tested the validity of PPP hypothesis have normally used a wide range of cointegration

[2] It is important to mention that there is some difference in the countries mentioned in Table 1: Indian top exporter partners and in Fig. 1: Time series plots of the India's real exchange rate with major trading partners. This is because CPI data was not available from where we sourced our data i.e., from the International Monetary Fund's International Financial Statistics CD-ROM. Therefore, we were unable to construct the real exchange rate series for the excluded countries.

[3] http://commerce.nic.in/publications/anualreport_chapter3-2009-10.asp.

techniques including the residual-based test for cointegration method proposed by Engle and Granger (1987), fully modified OLS procedure due to Phillips and Hansen (1990), Johansen (1988, 1991) multivariate cointegration technique, the autoregressive distributive lag (ARDL) approach postulated by Pesaran et al. (2001) and other tests with structural breaks in time series and panel framework or unit root test with and without incorporation of structural breaks.[4] However, in the present study we used a number of tests for validating the existence of PPP hypothesis. We used improved versions of the Dickey-Fuller type unit root tests, namely, the DF-GLS test of Elliot et al. (1996), as well as the nonlinear unit root of Kapetanios et al. (2003).

There are few studies investigating the validity of PPP theory in India but presenting conflicting empirical evidence on PPP hypothesis. So, the purpose of this study is to probe the long-run relationship between nominal exchange rate and relative prices for the India's major trading partners. The present empirical study contributes to this line of research by determining whether PPP theory holds for India's real exchange rate relative to a sample of her major trading partner countries (Hong Kong, Saudi Arabia, Singapore, UK and US), by taking into the account of asymmetric adjustment. The study uses data over the period of $1991M_1$–$2009M_2$ reflecting era after financial reforms were implemented in 1990s in all developing and emerging economies as well as in India. For this purpose, the TAR and M-TAR cointegration tests for unit root postulated by Enders and Granger (1998) are employed in this study. The major advantage of this approach is that it allows us to incorporate asymmetric adjustment in the cointegration process. We find that PPP does not hold for India's bilateral real exchange rate for major trading partners i.e., US, UK, Singapore, Saudi Arabia and Hong Kong. The TAR model was introduced by Tong (1978) and latter on modified by Tong and Lim (1980) and Tong (1983). The TAR and M-TAR cointegration tests are superior to traditional cointegration techniques. For example, the TAR models have limit cycles, amplitude dependent frequencies, and jump phenomena. These models are capable of producing asymmetric limit cycles and are suitable for time series data. The parameters of TAR models are computed simply and efficiently. At every stage, TAR estimators also have minimum variances by computing ordinary least square method. Last but not least, the TAR models have also been used successfully to explore asymmetries in macroeconomic variables over the course of the business cycle. These models have been widely used in the research for application due to lack of suitable data and failures to identify the threshold variables to estimate the threshold values.

This paper is organized as follows. Section 2 presents a brief review of literature followed by the data used in our study and methodology adopted for analysis in Sect. 3. Our empirical results are shown in Sects. 4 and 5 presents the conclusion of the paper.

2 A brief review of literature

There are some studies providing no evidence of PPP. For example, Layton and Stark (1990) for G-7 countries; Edison and Fisher (1991) for high income countries;

4 See the literature review part for related review.

Corbae and Ouliaris (1991) for Australia; Bahmani-Oskooee and Rhee (1992) and Kim (1995) for Korea; Bahmani-Oskooee (1995a) for 19 industrialized countries; Engle et al. (1997) for US cities; O'Connell (1998) for 59 developed and developing economies applying the Engle-Granger cointegration technique to the weighted-average price indices of G-7 countries, find little evidence of cointegration between the indices, thus rejecting PPP.[5]

However, some studies have found support for PPP. For example; Bahmani-Oskooee (1995b) by applying the Engle-Granger procedure to assess the Greek drachma versus the weighted price indices of 19 trade partners found evidences for PPP. Bahmani-Oskooee (1995c), after constructing nominal and real effective exchange rates for 22 LDCs based on each trading partner's 1985 import share, applies two stationarity tests and shows that PPP holds in 8 out of 22 countries.[6] Bahmani-Oskooee (1998), using the unit root test of Kwiatkowski, Phillips, Schmidt, Shin (KPSS) (Kwiatkowski et al. 1992a) finds support for PPP in 8 out of 11 Middle Eastern countries studied. Similar results are attained by Bahmani-Oskooee and Mirzaie (2000), who find support for PPP in most of the countries in their sample of 22 LDCs when the KPSS stationarity test is used. Similarly, Clemente et al. (1999) by applying various tests provided evidence of stationarity in the real effective exchange rates of 11 OECD countries however, only if up to two shifts in the mean are allowed. Thus the rate is able to revert to a moving mean. Cashin and McDermott (2003) show that slow reversion to parity-which is often interpreted by statistical tests as an outright lack of reversion-takes place in the real effective exchange rates of 20 industrial countries. Barlow (2003) exmianed the presence of purchasing power parity using data of Poland, Czech Republic and Romania. The empirical results found the valiadtion of PPP theory in Poland, Czech Republic but not in case of Romania. Bahmani-Oskooee and Kandil (2007a) validated PPP theory in 8 out of 14 MENA countries applying non-linear STAR framework.[7] Shabbir and Rashid (2008) have applied linear as well as nonlinear unit root test based on Taylor expansion to test for the presence of PPP hypothesis in the sought Asian countries.[8] Their results indicated that PPP exists in Pakistan, India, Sri Lanka, Bangladesh and China confirmed by nonlinear unit root tests.[9] Noman and Rahman (2010) applied exponential STAR framework to test the stationarity properties of exchange rates in South Asian countries and found that PPP hypothesis does not hold in India, Pakistan and Sri Lank and, results of about Bangladesh are inconclusive. Chang et al. (2010) applied threshold cointegration test developed by Enders and Dibooglu (2001) and concluded that PPP theory is valid in BRICs except China. In case of South Africa, Bonga–Bonga (2011) concluded that PPP hypothesis is weakly supported by VAR-X

[5] See for more details Cuddington and Hong (2000), Baum et al. (2001), Layton and Stark (1990).

[6] Shahbaz (2009), Shahbaz and Wahid (2010) did same exercise for Pakistan and Philippines respectively. Bahmani-Oskooee and Kutan (2008) investigated the impact of devaluation on domestic output of Central European economies.

[7] A detailed review of literature on PPP hypothesis is available in Bahmani-Oskooee and Hegerty (2009).

[8] Pakistan, India, Sri Lanka, Bangladesh and China.

[9] The results by Shabbir and Rashid (2008) may be biased because they did not mention which lag length criteria was used. It is well established fact that lag-length selection plays very crucial role in the outcome of test statistics.

approach. Cavusoglu and Telatar (2011) revisited PPP theory by using time-varying parameter approach and documented that PPP hypothesis exists in case of US. Baharumshah and Soon (2012) reported that currency crisis in Malaysia weakened the evidence of PPP hypothesis. Bozoklu and Kutlu (2012) utilized non-linear cointegration and reported that PPP is valid in Korea, Malaysia and Thailand but not in India, Indonesia, Mexico, the Philippines and South Africa following Haug and Basher (2011). Similarly, Boršič et al. (2012) used month frequency data by applying SURADF test developed by Breuer et al. (2002) and found that PPP theory is supported in some reforming European countries.

Note that all those adopted tests assumed a symmetrical relationship between exchange rate and relative prices by construction. In this conjunction, although there has been a widely-held belief of the symmetrical adjustment dynamics of nominal exchange rate towards the PPP equilibrium (see, for instance, Baum et al. 2001), there is no reason to pre-assume that the PPP equilibrium relationship, if any, must exist in a symmetrical fashion. In fact, it had been shown in Enders and Dibooglu (2001) and Liew (2004) that such adjustment process is asymmetry in nature. Liew (2004) asserted that the responses of market adjustment mechanism towards overvaluation and undervaluation of nominal exchange rates as compared to the PPP equilibrium follow asymmetric path. Besides, Enders and Dibooglu (2001) argued that asymmetric adjustment could be primarily due to prices that are sticky in the downward direction.

In addition, Enders and Granger (1998) demonstrated that conventional unit root and cointegration tests have low power in the presence of asymmetric adjustment. Furthermore, the above-mentioned tests are parametric in nature in which their results are dependent on the specification of the test. In sharp contrast, based on nonparametric tests, Liew et al. (2009) were able to provide evidence supportive of the PPP hypothesis for these countries. Therefore, it is possible that the unfavorable results from the transition economies are due to the negligence of asymmetric adjustment in the specification of the testing procedures. Though, there has recently been a growing consensus among the researchers that the real exchange rate exhibits nonlinearities, and hence, conventional unit root tests which are based on linearity assumption such as the Augmented Dickey Fuller (ADF) test and Phillips-Perron (PP) test and even more powerful tests like Kwiatkowski, Phillips, Schmidt, Shin (KPSS) (Kwiatkowski et al. 1992b) test and Ng and Perron (2001) test have low power in detecting the mean reversion process of exchange rate when nonlinearities are present. Reasons for the nonlinear adjustment are the presence of transactions costs that inhibit international goods arbitrage and official intervention in the foreign exchange market may be such that nominal exchange rate movements are asymmetric (see for example Taylor 2004; Juvenal and Taylor 2008; Reitz and Taylor 2008 among others). Kilian and Taylor (2003) also suggest that nonlinearity may arise from the heterogeneity of opinion in the foreign exchange market concerning the equilibrium level of the nominal exchange rate as the nominal rate takes on more extreme values, a great degree of consensus develops concerning the appropriate direction of exchange rate moves, and traders act as accordingly.[10]

[10] See Bozoklu and Kutlu (2012) for more details.

There are few studies which provided empirical evidence on the nonlinear adjustment of exchange rate. Examples in this group includes Michael et al. (1997), Baum et al. (2001), Parsley and Popper (2001), Taylor (2001), Taylor et al. (2001), Sarno et al. (2004), Chortareas and Kapetanios (2004), Erlat (2004), Hasan (2004), Liew et al. (2004), Shabbir and Rashid (2008), Wallace (2008), Bahmani-Oskooee and Gelan (2006), Chang et al. (2010), Haug and Basher (2011). A common feature of these studies is that they provide some support for nonlinear mean reversion in some real bilateral exchange rates. Nevertheless, it is important to mention that the findings of nonlinear adjustments does not necessarily imply nonlinear mean reversion i.e., nonlinear stationarity. Therefore, to test the stationary property of any data series which exhibits nonlinearity, stationarity tests based on a nonlinear framework must be applied and study make contribution to the literature by executing such test.

3 Data and methodology

Our empirical analysis covers a sample of five countries: Hong Kong, Saudi Arabia, Singapore, UK and US. For the analysis we employed monthly data and the time span is from January 1991 to February 2009. All consumer price indices, (CPI) (based on 2000 = 100), and nominal exchange rates of India relative to the trading partners data are taken from the International Monetary Fund's International Financial Statistics CD-ROM.[11] Results from two improved versions of Dickey-Fuller type unit root tests, namely, the DF-GLS test of Elliott et al. (1996) and Kapetanios et al. (2003)—a nonlinear unit root- are utilized. Elliott et al. (1996) proposed to extract the constant and trend effects from the series of interest using the general least squares (GLS) method, prior to the estimation of the Dickey and Fuller (1979) test, yielding the so-called DF-GLS test. It has been shown that the DF-GLS test has the best overall performance in terms of small-sample size and power, dominating the ordinary Dickey-Fuller test (Baum 2001; Vougas 2007). Vougas (2008) further demonstrated that the DF-GLS test has good size even when there is a neglected level or trend break under the null hypothesis. The DF-GLS test is applied on the real exchange rate series. Further, to cater for nonlinearity, Kapetanios et al. (2003) extended the DF and ADF unit root tests by allowing for nonlinear adjustment. Therefore, we used this test also in our analysis because conventional univariate unit root tests such as the ADF test have comparatively low power to reject a false null hypothesis of unit roots (see for example, Campbell and Perron 1991; Lothian and Taylor 1996, 1997) and are sensitive to the choice of lag length (see for example, Cuddington and Hong 2000). Kapetanios et al. (2003) proposed test is based on the following exponential smooth transition autoregressive (ESTAR) specification:

$$\Delta y_t = \gamma y_{t-1}\left[1 - \exp\left(-\theta y_{t-1}^2\right)\right] + \varepsilon_t \quad \theta \geq 0 \tag{1}$$

[11] The real exchange rate (RER) series of a country at time t is define as $(NER_t{*}CPIT_t)/CPII_t$, where *RERt* is the nominal exchange rate of India per foreign currency of trading partner, $CPIT_t$ and $CPII_t$ denote the consumer price indices of trading country and the India, respectively.

where y_t is the de-meaned or de-trended series of interest, ε_t is error with zero mean and constant variance, and $\left[1-\exp\left(-\theta y_{t-1}^2\right)\right]$ is the exponential transition function adopted in the test to present the nonlinear adjustment. The null hypothesis of a unit root in y_t (i.e. $\Delta y_t = \varepsilon_t$) implies that $\theta = 0$ $\left(\text{thus}\left[1-\exp\left(-\theta y_{t-1}^2\right)\right]=0\right)$. If θ is positive, it effectively determines the speed of mean reversion.

Kapetanios et al.'s (2003) test directly focuses on the θ parameter by testing the null hypothesis of non-stationarity H_0: $\theta = 0$ against the mean-reverting nonlinear alternative hypothesis H_a: $\theta > 0$. Because y_t in (1) is unidentified under the null, we cannot directly test H_0: $\theta = 0$. To deal with this issue, Kapetanios et al. (2003) reparameterize (1) by computing a first-order Taylor series approximation to specification (1) to obtain the auxiliary regression expressed by (2) below:

$$\Delta y_t = \delta y_{t-1}^3 + error \tag{2}$$

Assuming a more general case where the errors in (2) are serially correlated, regression (2) is extended to

$$\Delta y_t = \sum_{j=1}^{p} \rho_j y_{t-j} + \delta y_{t-1}^3 + error \tag{3}$$

with the p augmentations, which are used to correct for serially correlated errors. The null hypothesis of non-stationarity to be tested with either (2) or (3) is H_0: $\delta = 0$ against the alternative of H_1: $\delta < 0$. Kapetanios et al. (2003) show that the t-statistic for $\delta = 0$ against $\delta < 0$, i.e. t_{NL}, does not have an asymptotic standard normal distribution. They tabulate the asymptotic critical values of the t_{NL} statistics via stochastic simulations. In this paper, we estimate the t_{NL} statistics using regression (3) but for de-meaned and de-trended data series and refer to them as t_{NL1} and t_{NL2}, respectively. To obtain the de-meaned or de-trended data, we first regress each series on a constant or on both a constant and a time trend, respectively, and then we save the residuals. Thereafter, with saved residuals we estimated Eq. (3) and to avoid problem of serial correlation we choose appropriate lag-length. Further, to select the lag length (k), in order to avoid serial correlation, we use the 't-significance' approach[12] proposed by Hall (1994). This involves starting with a predetermined upper bound k. If the last included lag is significant, k is chosen. However, if k is insignificant,[13] it is reduced by one lag until the last lag becomes significant. If no lags are significant k is set equal to zero.

Finally, to cater for asymmetric adjustment, Enders and Granger (1998) generalized the Dickey-Fuller test to consider the null hypothesis of a unit root against the alternative hypothesis of a TAR model or M-TAR model. This Enders-Granger test can be specified as:

[12] The't-sig' approach has been shown to produce test statistics which have better properties in terms of size and power than information-based methods such as the Akaike Information Criterion or Schwartz Bayesian Criterion (see for example, Hall 1994, Ng and Perron, 1995).

[13] We used conventional level of significance that is 5% level of significance as a benchmark and fixed $k_{max} = 12$.

$$\Delta z_t = \rho_1 z_{t-1} I_t + \rho_2 z_{t-1}(1 - I_t) + \sum_{i=1}^{k} \alpha_i \Delta z_{t-i} + \varepsilon_t, \tag{4}$$

where $\varepsilon_t \sim i.i.d(0, \sigma^2)$; z_t is demeaned or/and detrended of real exchange rate, $y_t = \ln(q_t) + \ln(p_t^*) - \ln(p_t)$; q_t represents the nominal exchange rate defined as the price of foreign currency (that is trading partner) in domestic term; p_t^* and p_t are foreign and domestic price levels respectively. I_t is an indicator function that assumes the value of one if $\mu_{t-1} \geq \tau$, and zero if $\mu_{t-1} < \tau$ where $\mu_{t-1} = \{Z_{t-1}, \Delta Z_{t-1}\}$ and is the threshold value which governs the adjustment dynamic. Suppose $\mu_{t-1} < \tau$, the indicator function $I_t = 0$, such that

$$\Delta z_t = \rho_2 z_{t-1} + \sum_{i=1}^{k} \alpha_i \Delta z_{t-i} + \varepsilon_t, \tag{5}$$

and if $\mu_{t-1} \geq \tau$, $I_t = 1$so that

$$\Delta z_t = \rho_1 z_{t-1} + \sum_{i=1}^{k} \alpha_i \Delta z_{t-i} + \varepsilon_t, \tag{6}$$

Depending on specification of μ_{t-1}, the test is capable of detecting cointegration with TAR (when $\mu_{t-1} = Z_{t-1}$) and M-TAR (when $\mu_{t-1} = \Delta Z_{t-1}$) adjustments. Here the threshold,τ, is set endogenously using Chan's (1993) methodology.[14] The MTAR model allows the residual series to exhibit more momentum in one direction than the other. TAR and MTAR models allow the residuals to exhibit different degrees of autoregressive decay depending on the behavior of the lagged residual and its first difference, respectively.

Equation (4) encompasses two conventional augmented Dickey-Fuller (ADF) unit root tests, specified in Eqs. (5), (6). The principle of the Enders-Granger test is that if ρ_1 and ρ_2 are simultaneously equal to zero, the series is non-stationary (random walk). The null hypothesis of $\rho_1 = \rho_2 = 0$ may be tested by the $\Phi_\mu \, and \, \Phi_\mu^*$ test statistic for TAR and MTAR models respectively, which follows a nonstandard F distribution. The critical values for $\Phi_\mu \, and \, \Phi_\mu^*$ depend on the number of observations and the number of variables in the cointegrating vector. The critical values for the three-variable case have been tabulated by Enders and Dibooglu (2001). If the unit root hypothesis (i.e. null hypothesis of no cointegration) is rejected, the series is assumed to be stationary (mean-reverting), implying long-run relationship between exchange rate and relative price with asymmetric adjustment (i.e. hence alternative hypothesis of PPP holds good is accepted). If the null hypothesis of no cointegration is rejected, then one may proceed to test for the null

[14] Chan (1993) shows to obtain a consistent estimate of the threshold τ. For example, with the M-TAR adjustment mechanism, the consistent estimate of the threshold can be estimated by ordering the $\{\hat{\varepsilon}_t\}$ sequence in ascending order such that $\Delta\varepsilon_1^\tau < \Delta\varepsilon_2^\tau < \Delta\varepsilon_3^\tau < \ldots < \Delta\varepsilon_T^\tau$ where T denotes the number of usable observations. For each value of $\Delta\varepsilon_j^\tau$, the threshold $\tau = \Delta\varepsilon_j^\tau$ is set and the M-TAR model estimated in the form of Eqs. (5), (6). The estimated threshold yielding the lowest residual sum squares (RSS) is the consistent estimate of the threshold. To ensure an adequate number of observations in each regime, the standard procedure of using only the middle 80% of the observations as potential thresholds is followed.

hypothesis of symmetric adjustment (i.e.$\rho_1 = \rho_2$). If the null hypothesis of symmetric adjustment is rejected and $|\rho_1| > |\rho_2|$, then the model exhibits more decay for positive (changes in) errors.

4 Data analysis and findings

First of all we utilized a linear unit root test, an improvised version of DF test, DF-GLS test of Elliott et al. (1996), and present the results in Table 2.

Table 2 shows that none of the series is stationary based on this test, implying no evidence of long-run relationship between nominal exchange rates and relative prices of the five countries under consideration. Further we examined the PPP-hypothesis using nonlinear unit root test proposed by Kapetanios et al. (2003). Results of Kapetanios et al.'s (2003) test are reported in Table 3 for de-meaned and de-meaned and de-trended data series.

It is evident from the Table 3 that when we did analysis with de-meaned data series we find evidence for the support of PPP-hypothesis in case of US (at 1 % level of significance) and UK (at 5 % level of significance). However, for the de-meaned and de-trended data series we are able to reject the null hypothesis of non-stationrity and hence have support for holding of PPP-hypothesis for all five major trading partner countries of India. In the final step we have proceeded for estimation by incorporating

Table 2 Results of liner unit root analysis

Country	Constant	Constant and trend
United States	−0.6232 (0)	−1.0248(1)
Hong Kong	−0.4480 (0)	−0.5328 (1)
Saudi Arabia	−1.2222 (1)	−1.3875 (1)
Singapore	−0.8307 (0)	−1.0591 (0)
United Kingdom	−0.5739 (0)	−1.3377(1)

(1) Critical values −2.5756, −1.9422 and −1.6157 at 1, 5 and 10 % level of significance for constant and critical values are −3.4617, −2.9266 and −2.6340 at 1, 5 and 10 % level of significance for constant and trend model. (2) "k" denotes lag length and selection is based on SIC

Table 3 Results of nonlinear unit root analysis

Country	De-meaned (k)	De-meaned and De-trended (k)
United States	−3.5924 (0)***	−3.8237 (1)**
Hong Kong	0.4711 (12)	−3.8237 (1)**
Saudi Arabia	0.4711 (12)	−3.8237 (1)**
Singapore	−0.8469 (12)	−3.8237 (1)**
United Kingdom	−2.9441 (1)**	−3.8237 (1)**

(1) Critical values are −3.48, −2.93 and −2.66 for de-meaned data series. (2) Critical values are −3.93, −3.40 and −3.13 for de-meaned and de-trended data series at 1, 5 and 10 % level of significance respectively. (3) "k" denotes lag length and selection is based on SIC

the asymmetric adjustment in the cointegration process in the framework of TAR and MTAR models and results are reported in the following Table 4.

It is evident from Table 4 that null hypothesis of no cointegration is not rejected in any of the case in both specifications that is in case of TAR and MTAR models. This simply implies that PPP hypothesis do not hold in case of India with her major trading partners when consistent estimates of thresholds are allowed. Since, null

Table 4 Results of TAR and MTAR models analysis

Country	TAR [k]	MTAR [k]
United States		
Threshold	0.0928 [1]	−0.0022 [1]
ρ_1	−0.0218 (0.0342)	−0.0723 (0.0372)*
ρ_2	−0.0800 (0.0218)***	−0.0602 (0.0213)***
Φ_μ *and* Φ_μ^*	6.9213	5.8802
$t(\rho_1 = \rho_2)$	2.0535 {0.1533}	0.0794 {0.7782}
Q(3)	1.6336 {0.6517}	1.7335 (0.6295}
Hong Kong		
Threshold	0.0615[1]	−0.0095 [1]
ρ_1	−0.0096 (0.0058)*	0.0040 (0.0064)
ρ_2	0.0037 (0.0029)	0.0003 (0.0029)
Φ_μ *and* Φ_μ^*	2.1634	0.2013
$t(\rho_1 = \rho_2)$	4.1933{0.0418}	0.2715 {0.6028}
Q(3)	2.7862 {0.4257}	4.1280 {0.2479}
Saudi Arabia		
Threshold	−0.0249 [1]	−0.0094 [1]
ρ_1	−0.0096 (0.0034)***	−0.0012 (0.0053)
ρ_2	−0.0015 (0.0028)	−0.0056 (0.0024)**
Φ_μ *and* Φ_μ^*	3.9472	2.6156
$t(\rho_1 = \rho_2)$	3.1724 {0.0763}	0.5662 {0.4525}
Q(3)	5.5799 {0.1339}	7.0638 {0.0698}
Singapore		
Threshold	−0.0747[1]	−0.0018 [1]
ρ_1	−0.0164 (0.0066)**	−0.0011 (0.0124)
ρ_2	−0.0003 (0.0060)	−0.0086 (0.0047)*
Φ_μ *and* Φ_μ^*	3.0614	1.6264
$t(\rho_1 = \rho_2)$	3.1435 {0.0776}	0.3127 {0.5765}
Q(3)	0.5634 {0.9047}	1.0739 {0.7833}
United Kingdom		
Threshold	0.1371[1]	0.0223[1]
ρ_1	−0.0385 (0.0137)***	−0.0634 (0.0253)**
ρ_2	0.0040 (0.0158)	−0.0117 (0.0112)
Φ_μ *and* Φ_μ^*	3.9682	3.6939
$t(\rho_1 = \rho_2)$	4.0192 {0.0462}	3.4803 {0.0634}
Q(3)	2.3191 {0.5088}	3.2035 {0.3613)

(1) "k" denotes lag-length which is the number of auxiliary regressors in the TAR and MTAR model and based on SIC and AIC. (2) Φ_μ *and* Φ_μ^* are the F-statistics for the null hypothesis of no-cointegration or $(\rho_1 = \rho_2 = 0)$under the TAR and MTAR specifications respectively. (3) The 5 % critical values of Φ_μ is 7.30 for the TAR model and for Φ_μ^* are 7.96 for the MTAR model (4) $t(\rho_1 = \rho_2)$is the t-statistics for the test of symmetry. (4) Q(3) is the F-statistics for serial correlation of third degree. (5) *, ** and *** denotes significance at 10, 5 and 1 % level of significance respectively. (6) in () and { } we mentioned standard errors and *P* values respectively

Source: Author's calculation

hypothesis of no cointegration is not rejected in any of the case mentioned above mentioned therefore, no need to proceed to test for the null hypothesis of symmetry (i.e. $t(\rho_1 = \rho_2)$) under the TAR and MTAR specifications. However, the most interesting case is when we used linear and nonlinear unit root test of time series we find, in general, that PPP hypothesis holds however, when we used TAR and MTAR models by allowing for consistent estimates of the threshold we are unable to find even single evidence in support of PPP hypothesis. This shows that exchange rate is not monetary phenomenon as Indian economy is following flexible exchange arte system.

5 Conclusions

This study has made an attempt to test the validity of PPP hypothesis for India by using data of real exchange rate in bilateral framework. Study utilized data which spans from January 1991 to February 2009, monthly observations, and countries included are: Hong Kong, Saudi Arabia, Singapore, U.K and U.S. For analysis study used improved versions of Dickey-Fuller type unit root tests, namely, the DF-GLS test of Elliott et al. (1996), as well as the nonlinear unit root of Kapetanios et al. (2003). Finally, to cater for asymmetric adjustment, Enders and Granger (1998) generalized the Dickey-Fuller test to consider the null hypothesis of a unit root against the alternative hypothesis of a TAR model or M-TAR model which is also utilized in our analysis.

Study finds by using DF-GLS test that in none of the case, null hypothesis is rejected however, when nonlinear unit root test proposed by Kapetanios et al. (2003) is utilized, study finds contrasting results. That is study rejected the null hypothesis of non-stationrity and hence finds support for holding of PPP hypothesis for all five major trading partner countries of India. Further, when TAR and MTAR models are analyzed null hypothesis of no cointegration is not rejected even for single case hence provides no evidence for the support of PPP hypothesis in case of India with her major trading partners. Hence, we conclude that in case of India we do not have evidence to support for PPP hypothesis with major trading partners. This reveals that intermediate goods face high barriers to trade among these sample countries. This supports the argument that Indian government has not been able to strike out the proper balance between flexibility and stability between real bilateral exchange rates and thus unable to maintain the confidence in the domestic currency that has been evident from the recent fall of rupee in relation to the US $. Overall our results suggest that India needs to design more effectively the bilateral trade relationship. Further, from the results of the sample of five major trading nations we found that overall stability of the real effective exchange rate can be achieved only under the fixed exchange rate system. This is because a fixed exchange rate policy provides stability and confidence to hold the Indian currency nonetheless; there is risk of realizing an overvalued real exchange rate that adversely affects competitiveness and export growth of India. Similarly as Bahmani-Oskooee and Kandil (2007b) recommended that an ‘economy that is highly dependent on imported goods should opt for a fixed exchange rate policy to stabilize inflationary expectations and insulate the economy from nominal shocks whereas a flexible exchange rate policy is better for an economy which faces real shocks as flexible exchange rate policy has flexibility to absorb real shocks and minimize their adverse effects on the economy. Further, if Indian government is not able to minimize the deviations around the targeted equilibrium, she may face aggravating adverse effects on output growth and economic stability in the short-run. This is because shocks to the exchange rate may create positive or negative effects on output growth depending on supply and demand channels and also on the economic condition of a nation in consideration.

In future, present study can be augmented to revisit PPP hypothesis in India particularly and in African countries, Latin American countries, European Countries and Asian countries by applying structural break (single and two structural breaks) unit root tests and structural break cointegration approaches following Baharumshah and Soon (2012). The structural break unit root tests are Perron test by Perron (1989); Z–A test by Zivot and Andrews (1992); LP test by Lumsdaine and Papell (1997); Clemente-Montanes-Reyes test by Clemente et al. (1998); SL test by Saikkonen and Lütkepohl (2002); LS test by Lee and Strazicich (2003, 2004); NP test by Narayan and Pop (2010) and many more can be applied. Similarly, cointegration with single unknown structural break developed by Carrion-i-Silvestre and Sanso (2006) and with two unknown structural breaks by Hatemi-J (2008) can also applied to test the existence of cointegration for long run relationship between the variables.

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