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Nonlinear adjustment effects in the purchasing power parity

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ABSTRACT: This study examines nonlinear adjustment effects in the purchasing power parity (PPP) between South Africa and her main currency trading partners; namely, the US, the UK, the Euro area, China and Japan. We use monthly data of the nominal exchange rates and domestic price level data collected between the periods 1971-2014. The empirical study is conducted using nonlinear unit root and asymmetric cointegration analysis. Our empirical results show significant asymmetric PPP effects between South Africa and her main trading partners with causal effects flowing from exchange rates to price differentials.

Keywords: Purchasing power parity (PPP); Threshold co-integration, Momentum threshold autoregressive (MTAR) model; Threshold Error correction (TEC) model; South Africa.

JEL Classification Code: B22, C22, C32, E31, E58, F31.

1 INTRODUCTION

Up to date, the purchasing power parity (PPP) represents one of the oldest and yet remains one of the most controversial doctrines existing within the economic paradigm. The underlying notion of the PPP hypothesis presents deviations from the parity as profitable commodity arbitrage opportunities which, if exploited, will tend to bring the exchange rate towards parity (Brissimis et. al., 2005). Alternatively stated, the PPP relationship predicts a constant equilibrium level at which exchange rates converge in such a way that foreign currencies possess the same purchasing power and any change in the exchange rate between two countries' currencies is determined by the relative price ratio between the domestic and foreign countries (Azail et. al., 2001). Therefore, much of open or external macroeconomy policy is based on the PPP hypothesis and estimates of the PPP are frequently used in determining the degree of misalignment of the nominal exchange rate; evaluating the appropriate policy response to detected misalignments in the exchange rate; determining the setting of exchange parities, and international comparison of national income levels (Taylor and Taylor, 2004).

From an empirical perspective, the PPP hypothesis requires a real exchange rate to either evolve constantly over time or at least exhibit mean reverting behaviour with no stochastic trend (Bozoklu and Kutlu, 2012). The empirical validation of stationarity in real exchange rates is important because an exchange rate characterized by a unit root will ensure that the time series will not revert back to it's steady state equilibrium in the face of a macroeconomic shock. Unit root behaviour in the real exchange rate will thus impact the forecastability of exchange rates since a non-stationary exchange rate particularly implies that the best forecast of the following year's exchange rate is the most recently observed exchange rate values and thus the predictability of such a time series never tends to an average value. This, in turn, will undermine policymakers ability to discern whether exchange rates are overvalued or undervalued, more especially if a monetary approach is used in determining the exchange rate level. Ultimately, a stationary exchange rate process is most desirable for open or external macroeconomic policy, especially in emerging or developing economies where exports are a principal source of economic growth and monetary authorities are typically concerned about the unpredictability of exchange rates which are likely to affect net exports as well as the cost of servicing foreign-currency-denominated-debt.

There exists an almost unanimous consensus stating that that the PPP is not a short-run phenomenon of exchange rate movements and this has been so since monetary policy shifts to floating exchange rates regimes as experienced by a considerable number of central banks worldwide in the late 1990's. Since then, the PPP hypothesis has been modelled as a long-run parity condition linking relative prices and the exchange rates. However, even in modelling long-run movements in real exchange rates, such attempts by researchers have typically been met with mixed results, more prominently for African economies. This phenomenon is conveniently iterated in a study conducted by Liu and Su (2011). One plausible explanation for the inconclusiveness found in previous studies is the failure of these studies to account for a possible nonlinear adjustment in the PPP relationship. Nakagawa (2010) argue that nonlinearity may arise in the presence of the transactions costs that preclude goods-market arbitrage; and only when the price differentials become large enough to outweigh the costs, will arbitrage operate to eliminate deviations from PPP. Also Bozoklu and Kutlu (2012) claim that the disparity of price indices, the existence of non-tradable goods, trade barriers and imperfect competitive market structures also contribute towards invalidating the assumption of a linear PPP hypothesis in the long-run. Moreover, Holmes and Wang (2006) attribute asymmetric behaviour in exchange rates to the reluctancy commonly shown by Central Banks in facilitating depreciation of the nominal exchange rate in a regime of managed floating as well as to heterogeneity of participants in the foreign exchange market in terms of agents expectations formation or investors objectives.

In screening through the former evidence as presented in previous case studies, one is able to observe that there generally exist two strands of literature which empirically examine the significance of asymmetries in the PPP relationship. The first strand of these studies examines the asymmetries in the PPP hypothesis by examining the integration properties of a series of real exchange rates through the use of asymmetric unit root tests (Kim and Moh, 2010; Yoon, 2010; Liu and Su, 2011). The second strand of studies applies asymmetric co-integration techniques in examining the correlation between real exchange rates and differences in the price indices (Baum et. al. 2001; Holmes and Wang, 2006; Nakagawa, 2010). Generally, research academics have, for a variety of empirical or methodological rationale, preferred one approach over the other but rarely do economists opt to examine or use both approaches simultaneously. Provoked by this, our study contributes to the existing literature by applying asymmetric unit root tests and threshold co-integration analysis for the PPP hypothesis concerning South African economy relative to her trading currency partners. These trading partners are the United States (US); the United Kingdom (UK); the Euro area; China and Japan. Moreover, our study examines causality effects between the relative price differentials and exchange rates between South Africa and these aforementioned trading partners. We consider this research as a worthwhile undertaking since such an empirical exercise, to the best of our current knowledge, has not been conducted for South Africa relative to her main trading currency partners.

Having presented a motivation for the study, we present the remainder of the paper as follows. Section two provides an outline of how to test the PPP hypothesis using asymmetric unit root tests of Kapetanois and Shin (2006). Section three then outlines the momentum threshold autoregressive (MTAR) and threshold error correction (TEC) model of Enders and Silkos (1998) used to examine threshold co-integration effects in the PPP hypothesis. Section four introduces the empirical data and proceeds to conduct the empirical analysis on the empirical data. Section five of the paper concludes with policy implications as obtained from the empirical analysis.

2 PPP AND UNIT ROOT TESTS

According to Darby (1980), Haikko (1992) and Taylor and Taylor (2004) there are two distinct versions under which the PPP hypothesis might hold. Firstly, there is the absolute version of the PPP hypothesis which strictly adheres to the "law of one price" within an integrated and competitive market; and assumes homogeneity and substitutability of the goods with no transaction costs, tariffs, quotas and other trade barriers (Kargbo, 2004). The absolute PPP theory can be viewed as modification of the quantity theory of money in an open economy, in which an increase in the supply of money leads to a simultaneous increase in the price level and a decline in the exchange rate (Haikko, 1992). This relation can be captured using the following functional form:

$$\mathcal{P}_t^d = \epsilon_t \mathcal{P}_t^f \tag{1}$$

Where ϵ_t is the nominal exchange rate defined as the unit price of foreign currency in terms of home currency; whereas $\mathcal{P}_t^d$ and $\mathcal{P}_t^f$ are the local and foreign price levels, respectively. However, in practice, the absolute version of the PPP hypothesis has generally failed for three main reasons. Firstly, the absolute PPP theory seemingly holds only when the purchasing power of a unit for currency is exactly equal in both the domestic economy and the foreign economy (Taylor and Taylor, 2004). In other words, the absolute theory is strictly dependent upon the law of one price, which has been proved not to hold – even on average (McChesney et.al., 2004). Secondly, price levels in different countries are computed using imperfect price indexes and, as a result, the simple ratio of the price levels may not be an adequate measure of the equilibrium exchange rates (Haikko, 1992). Thirdly, deviations from absolute PPP may occur on account of transport costs, tariffs and differential speeds of adjustments in the goods and foreign exchange markets, of which the absolute PPP hypothesis does not take into consideration (Shirley, 2013).

Due to the aforementioned arguments, most economists and research academics have almost exclusively turned to their attention towards the use of the second version of the PPP hypothesis; namely, the weak or relative version of the PPP hypothesis. Generally, the relative version of the PPP hypothesis is favoured as a more effective measure of the equilibrium exchange rate since it follows directly from the absolute PPP and the relative PPP hypothesis may also hold when the absolute PPP fails to hold. Pragmatically, the weak or relative version of the PPP hypothesis casts the theory in terms of changes in relative prices and the exchange rates and consequentially, researchers commonly opt to use a logarithmic version of the PPP hypothesis as specified below:

$$log\epsilon_t = \alpha + \beta_1 log\mathcal{P}_t^d - \beta_2 log\mathcal{P}_t^f + \mu_t \qquad (2)$$

By further defining $log\pi_t^* = log\mathcal{P}_t^d - log\mathcal{P}_t^f$, one can re-specify equation (2) as a restricted form of the relative version of the PPP hypothesis as follows:

$$log\epsilon_t = \alpha + \beta log\pi_t^* + \mu_t \qquad (3)$$

Bahmani-Oskooee and Gelan (2006) note that under a floating exchange rate system, such as that adopted by the South African Reserve Bank (SARB), a country's nominal exchange rate may depreciate against one currency and appreciate against another. This renders it more feasible to rely on the real effective exchange rate for purposes of examining unit roots in the time series. Consequentially, researchers typically extend equation (2) to incorporate the real exchange rate in determining the equilibrium level of exchange rates and as a result, rely on the real exchange rate, as opposed to the nominal exchange rate, in validating the PPP hypothesis under the implementation of specified unit root tests. By definition, the real exchange rate is the nominal exchange rate (i.e. domestic price of foreign currency) multiplied by the ratio of national prices (i.e. domestic price level divided by foreign price level); and thus provides a measure of the purchasing power of a unit of foreign currency in the foreign currency relative to the purchasing power of an equivalent unit of domestic currency in the domestic economy (Taylor and Taylor, 2004). By denoting τ_t as the real exchange rate, we can substitute the real exchange rate formulae (i.e.$\frac{\tau_t}{\epsilon_t} = \frac{\mathcal{P}_t^d}{\mathcal{P}_t^f}$) into equation (2). In further re-arranging the terms and further converting the variables into logarithmic form, we can obtain the following PPP regression equation:

$$log\tau_t = log\epsilon_t - \beta log\pi_t^* \qquad (4)$$

From the equation (4), the real exchange rate, τ_t, may be, for empirical purposes, interpreted as a measure of deviation from PPP equilibrium. Thus, in order to validate the PPP hypothesis, one can simply examine the integration properties of the real exchange rate. Typically, testing for stationarity involves placing the real exchange rate subject to the following generalized autoregressive (Dickey-Fuller-type) regression:

$$\tau_t = \phi\tau_{t-1} + \xi_t \qquad (5)$$

Where ϕ is the least square estimate and $\xi_t$is the associated normally distributed error term. For the PPP hypothesis to be valid, the stationary hypothesis of $|\phi| < 1$ should not be capable of being rejected such that the evolution of the real exchange rate is stationary. This implies that the time series a stable, mean reverting autoregressive process such that deviations from the PPP are only temporary. However, the assumption of a linear data generating process may be trivializing the issue. If indeed the real exchange rate evolves as a nonlinear process, then linear unit root tests will have very low power to reject the null hypothesis of a unit root against the alternative of a unit root. Consequentially the observed time series variables may alternatively be tested for unit roots using appropriate nonlinear unit root testing procedures. This argument is iterated in Taylor (2001) and Bec et. al. (2004) who find that nonlinear unit root testing procedures on the real exchange rate for European economies produces superior testing power in comparison to those obtained for the conventional Dickey-Fuller tests. Taking note of this, our study relies on the nonlinear unit root testing procedure of Kapetanois and Shin (2006) which is based on the following three-regime threshold autoregressive (TAR) auxiliary regression:

$$\Delta Y = X(\gamma)\phi + \mathrm{v} \qquad (6)$$

Where:

$$\phi = (\phi_1, \phi_2)'; \;\; \Delta Y = \begin{pmatrix} \Delta\tau_1 \\ \Delta\tau_2 \\ \vdots \\ \Delta\tau_T \end{pmatrix}; \;\; X(\tau) = \begin{pmatrix} \tau_0(\gamma_1) & \tau_0(\gamma_2) \\ \tau_1(\gamma_1) & \tau_1(\gamma_2) \\ \vdots & \vdots \\ \tau_{T-1}(\gamma_1) & \tau_{T-1}(\gamma_2) \end{pmatrix}; \text{ and } \; v = \begin{pmatrix} \xi_1 \\ \xi_2 \\ \vdots \\ \xi_T \end{pmatrix}$$

Whereby γ_1 and γ_2 denoting the first and second threshold estimates, respectively, and the threshold functions for the first and last regimes (third regimes) are given by $I.\{\epsilon_i \leq \gamma_1\}$ and$I.\{\epsilon_i > \gamma_2\}$, respectively. From the aforementioned, the joint null hypothesis of a linear unit root (i.e. $H_0: \phi_1 = \phi_2 = 0$) can be tested against the alternative of a three regime stationary process with a unit root process existing in the middle regime (i.e. $H_1: \phi_1, \phi_2 < 0$) and these hypotheses can be tested using a standard Wald statistic computed as:

$$\mathcal{W}_{\gamma_1,\gamma_2} = \hat{\phi}'[Var(\hat{\phi})]^{-1}\hat{\phi} \tag{7}$$

Where $\hat{\phi}$ is the ordinary least squares (OLS) estimate of ϕ. However, given that the threshold parameters are unknown a prior, Kapetanois and Shin (2006) consider three commonly used summary statistics based on the supremum, average and exponential average variations of the Wald statistic as defined below in equations (8) through (10).

$$\mathcal{W}_{sup} = \sup_{(i\varepsilon\Gamma)} \mathcal{W}_{(\gamma_1,\gamma_2)} \tag{8}$$

$$\mathcal{W}_{ave} = {}^{1}\!/_{\#\Gamma} \sum_{i=1}^{\#} \mathcal{W}_{(\gamma_1,\gamma_2)} \tag{9}$$

$$\mathcal{W}_{exp} = {}^{1}/_{\#\Gamma} \sum_{i=1}^{\#} {}^{\mathcal{W}_{(\gamma_1,\gamma_2)}}/_{2} \qquad (10)$$

The optimal values of the threshold estimates, γ_1 and γ_2, are obtained by maximizing the Wald statistics over a search grid and then constructing summary statistics for the obtained threshold estimates. In the spirit of Kapetanois and Shin (2006), we employ the aforementioned nonlinear unit root testing procedure to three empirical cases, namely; (i) the case of a zero mean process; (ii) the case of a process containing a non-zero mean; and (iii) the case of a process containing both non-zero mean and linear trend. The associated asymptotic distributions are therefore computed using a de-meaned and the de-trended standard Brownian motion in the construction of the associated Wald test statistic.

3 PPP AND CO-INTEGRATION ANALYSIS

The equilibrium relationship captured by the absolute version of the PPP (as an aggregate interpretation of the law of one price) assumes that perfect commodity arbitrage acts an error correction mechanism to force the Rand price of a consumption bundle of South African goods in line with the Rand price of a common bundle of foreign goods. Since a cointegrated system allows individual time series to be integrated of order one, but requires a linear combination of the series to be stationary, PPP is testable using the theory of co-integration processes (Corbae and Ouliaris, 1988). From a co-integration perspective, the PPP doctrine suggests that nominal exchange rates should be determined according to the differences between foreign and domestic exchanges rates of inflation (Ozdemir, 2008). In this regard, a number of empirical studies are concerned with testing the PPP by examining whether nominal exchange rates, ϵ_t, and the differences between domestic and foreign price levels, π_t^* are cointegrated, that is, whether these time series variables move together over time. This can be empirically achieved by re-arranging equation (3), to resemble the Engle-Granger co-integration theorem for the PPP hypothesis and can be expressed as follows:

$$\mu_t = log\epsilon_t - log\pi_t^* \quad (11)$$

From regression (11), non-spurious co-integration effects or validity of the PPP hypothesis is assumed to exist under the integration conditions $log\epsilon_t$~I(1), $log\pi_t^*$~I(1) and μ_t~I(0); such that the nominal exchange rates and the differences in price indexes should increase monotonically over time with μ_t being the stationary equilibrium error of the co-integration relation. Therefore, the standard Engle-Granger procedure for ensuring co-integration between a pair of time series variables involves testing as to whether the equilibrium error, μ_t, is a mean reverting process. However, as previously mentioned, the relation between exchange rates and national price levels can, in reality, be affected by several factors including transport and information costs, imperfect competition, technological changes, factor supplies trade restrictions and non-traded goods and services. Kargbo (2003) also argues that changes in the monetary policy regimes as well as financial liberalization and losing of restrictions on capital inflows over the last two decades or so may be further account for rationally assuming nonlinearity in adjustment equilibrium process between aggregate prices and exchanges. Empirically, Cheung and Lai (1993) propose that the imposition of symmetry and proportionality conditions in analysing the PPP co-integration relationship can cause the restricted models to ignore possible interactions in the determination of exchange rates prices that are permitted in the unrestricted model. Furthermore, a number of econometricians such as Blake and Fomby (1997), Hansen and Seo (2002) and Seo (2006) have all demonstrated that linear co-integration tests may fail if the equilibrium adjustment process for the time series is indeed asymmetric. Therefore it is possible that linear adjustment leads to poor results of the equilibrium relationship because conventional co-integration tests do not take into account asymmetric equilibrium adjustment. In summing it up, the aforementioned arguments depict that models of exchange rate determination may depict fundamental differences in speeds of adjustment between exchange rates and price levels. Therefore, in line with Enders and Silkos (2001), we deviate from the assumption of linear co-integration and model the equilibrium error term (i.e. ξ_{t-1}) as the follows:

$$\varDelta \mu_t = I_t \rho_1 \xi_{t-1} + (1 - I_t) \rho_2 \xi_{t-1} + \sum_{i=1}^{p} \beta_i \, \Delta \xi_{t-i} + \varepsilon_t \qquad (12)$$

And thereafter apply the following co-integration tests for (i) stationarity of the equilibrium error term (ii) normal co-integration effects; and (iii) asymmetric co-integration effects, which are respectively implemented under the following hypothesis:

$$H_0^{(1)} : \rho_1, \rho_2 < 0 \qquad (13)$$

$$H_0^{(2)} : \ \rho_1 = \rho_2 = 0 \qquad (14)$$

$$H_0^{(3)} : \rho_1 = \rho_2 \qquad (15)$$

The threshold co-integration regression as specified in equation (12), can assume two primary functional forms. The first is a standard threshold autoregressive (TAR) form which is dictated by the following indicator functions for a zero threshold level and a consistent threshold estimate (c-TAR) specifications which are respectively denoted as:

$$I_{.t} = \begin{cases} 1, if\, \xi_{t-1} \geq 0 \\ 0, if\, \xi_{t-1} < 0 \end{cases} \qquad I_{.t} = \begin{cases} 1, if\, \xi_{t-1} \geq q \\ 0, if\, \xi_{t-1} < q \end{cases} \tag{16}$$

The second functional form for the threshold regression is given by a momentum threshold autoregressive model (MTAR) which differs from the standard TAR specifications since it captures large and smooth changes or capture spiky adjustments in the co-integration equilibrium relationship in a series whereas the TAR model is designed to whereas the TAR model is limited to capturing the depth of movements in the equilibrium residuals. The indicator functions for the MTAR with a zero threshold and the MTAR model with a consistent threshold estimate (c-MTAR) are respectively specified as:

$$M_{.t} = \begin{cases} 1, if\, \Delta\xi_{t-1} \geq 0 \\ 0, if\, \Delta\xi_{t-1} < 0 \end{cases} \qquad M_{.t} = \begin{cases} 1, if\, \Delta\xi_{t-1} \geq q \\ 0, if\, \Delta\xi_{t-1} < q \end{cases} \tag{17}$$

Where Δ denotes a first difference operator. Since the threshold variable under the c-TAR and c-MTAR models (i.e. q), are unknown a prior, the threshold co-integration regression (12) is estimated by ordering the threshold variable, q, in ascending order such that $q_0 < q_1 < \cdots < q_T$,where T is the number of observations used after truncating the upper and lower 15 percent of the observations. In accordance with Hansen (2000), the true threshold estimates is one which minimizes the residual sum of squares of the estimated regression equations.

According to the granger representation theorem, an error correction model can be estimated once a pair of time series variables is found to be cointegrated. When the presence of threshold co-integration is validated, the error correction model can be modified to take into account asymmetries as is demonstrated in Blake and Fombly (1997) and Enders and Silkos (2001). The asymmetric error-correction model also can exist between a pair of time series variables of $log\epsilon_t$ and $log\pi_t^*$ when they are formed in an asymmetric co-integration relationship. The TAR-threshold error correction (i.e. TAR-TEC) model can be expressed as:

$$\Delta\epsilon_t = \lambda_{11} I_t \xi_{t-1} + \lambda_{12}(1 - I_t)\xi_{t-1} + \sum_{i=1}^{p} \varphi_{i1}\Delta\epsilon_{t-1} + \sum_{i=1}^{p} \psi_{i1}\Delta\pi_{t-1}^* + \mathrm{v}_{t1} \qquad (18)$$

$$\Delta\pi_t^* = \lambda_{21} I_t \xi_{t-1} + \lambda_{22}(1 - I_t)\xi_{t-1} + \sum_{i=1}^{p} \varphi_{i2}\Delta\epsilon_{t-1} + \sum_{i=1}^{p} \psi_{i2}\Delta\pi_{t-1}^* + \mathrm{v}_{t2} \qquad (19)$$

Whereas the MTAR-threshold error correction (i.e. MTAR-TEC model is specified as:

$$\Delta\epsilon_t = \lambda_{11} M_t \xi_{t-1} + \lambda_{12}(1 - M_t)\xi_{t-1} + \sum_{i=1}^{p} \varphi_{i1}\Delta\epsilon_{t-1} + \sum_{i=1}^{p} \psi_{i1}\Delta\pi_{t-1}^* + \mathrm{v}_{t1} \qquad (20)$$

$$\Delta\pi_t^* = \lambda_{21} M_t \xi_{t-1} + \lambda_{22}(1 - M_t)\xi_{t-1} + \sum_{i=1}^{p} \varphi_{i2}\Delta\epsilon_{t-1} + \sum_{i=1}^{p} \psi_{i2}\Delta\pi_{t-1}^* + \mathrm{v}_{t2} \qquad (21)$$

The indicator functions as given in regressions (16) and (17) are respectively applied for the TAR-TEC and MTAR-TEC specifications. Through the above described systems of error correction models, the presence of asymmetries between the variables could initially be tested by examining the signs on the coefficients of the error correction terms; whereas granger causality tests can be implemented by using a standard F-test to examine whether the regression coefficients from the error correction models are significantly different from zero. Pragmatically, the null hypothesis of no error correction mechanism can be tested as:

$$H_0^{(4)}: \lambda^+ \xi_{t-1}^+ = \lambda^- \xi_{t-1}^- \qquad (22)$$

Whereas, the null hypothesis that the price differentials do not lead to nominal exchange rate is tested as:

$$H_0^{(5)}: \alpha_k = 0;\ i = 1, \dots, k \qquad (23)$$

And the null hypothesis that the nominal exchange rate does not lead to changes in price differentials do not lead to nominal exchange rate is tested as:

$$H_0^{(6)}: \beta_k\ = 0;\ i = 1, \dots, k \qquad (24)$$

4 EMPIRICAL ANALYSIS

In this section of the paper, we present the data description, the unit toot tests results as well as the MTAR empirical estimates and the causality analysis.

4.1 *DATA DESCRIPTION*

Our data set comprises of a total of 190 monthly observations collected between the periods of January 1971 to December 2014. For empirical purposes, it would have been more desirable to employ a longer span of data, but due to data availability constraints, consistent monthly data could only be collected from the period of 1998 onwards. The data used in our empirical analysis comprises of the nominal foreign exchange rate and price indices for South Africa and her main exchange currency partners. In particular, the collected price series are based on the total consumer price index (CPI) for South Africa, the United States (US), the United Kingdom (UK), the Euro area and China. Similarly, the nominal exchange rates are based on the nominal value of the Rand against the currencies of her main exchange partners namely against the US dollar (i.e. $\tau_t/us\$$); the British pound (i.e. $\tau_t/uk£$), the Euro (i.e. $\tau_t/euro€$), the Chinese Renminbi (i.e. $\tau_t/china¥$) and the Japanese Yen (i.e.$\tau_t/japan¥$). As a point of convenience as well as consistency, all price indices are collected from the International Monetary Fund (IMF) International Financial Statistics (IFS) database whereas the remainder of the data (i.e. the nominal exchange rates) is collected from the South African Reserve Bank (SARB) database. Finally, in line with Frankel and Rose (1996) as well as Akinboade and Makina (2006), we construct the domestic-based real exchange rate against all the other currency partners, using the relative form of the PPP hypothesis as previously specified in regression equation (4) (i.e. $log\tau_t = log\epsilon_t - \beta log\pi_t^*$). Furthermore, all empirical time series data is transformed into logarithmic form a prior.

A number of noteworthy stylized facts associated with the empirical data provide motivation for the use of asymmetric econometric techniques in analysing the PPP relationship between South Africa and her main currency trading partners. For instance, it is worth noting that empirical data covers an era in which most Central Banks worldwide experienced significant shifts in their conduct of monetary policy. Most notable of these monetary policy shifts are the adoption of an official inflation targeting regime as pursued by the SARB in 2002; independence of monetary policy in the UK in 1998; the Bank of Japan's adoption of a zero interest rate policy in 1998; and China's shift to a more "prudent" monetary policy in 2011. These is an important observation since these policy shifts further motivate the need to account for asymmetries in the empirical analysis of PPP behaviour between South Africa and her main currency exchange partners.

4.2 *UNIT ROOT TESTS*

Having put our data collection and formation into perspective; attention can now be turned towards examining the integration properties of the individual time series under observation. Even though the sole verification of stationarity in the real exchange rate is necessary in directly assessing the validity of the PPP hypothesis, we also extend our unit root tests towards the nominal exchange rates and the differences in the price indices as a preliminary step towards the co-integration analysis. As previously mentioned we perform the unit root tests with a zero-mean process, with an intercept and also with a trend and an intercept. We select the number of lags of the unit root tests based on the general-to-specific rule and decide on the optimal lag length as the system which produces the lowest Alkaike information criterion (AIC) decision rule. Table 1 below present the empirical results of the Kapetanois and Shin (2006) unit root tests as employed on the time series variables.

{INSERT TABLE 1 HERE}

In referring to the results reported in Table 1, we are able to reject the null hypothesis of a unit root in favour of a stationary three-regime TAR process for all observed time series when the unit root test is performed using the supremum and the exponential average on the Wald statistics. The evidence is less conclusive when the average on the Wald statistic is used to evaluate the integration properties of the time series variables. Generally these results provide us with preliminary evidence of PPP behaviour between South Africa and her main trading partners. One of the most interesting or noteworthy aspects of Kapetanois and Shin (2006) unit root tests is that they render the time series variables as a regime-switching processes consisting of both a stationary part as well as a unit root part. This is important towards our empirical analysis, since on one hand, this can render the stationary part of the computed real effective exchange rate as being in compliance with the PPP hypothesis, and on the other hand, it renders the unit root portion of the nominal exchange rate times series and the price differential time series as being providing preliminary evidence of PPP in the cointegration sense. To illustrate this point, consider the threshold estimates which determine the rand value at which the real exchange rate is found to be stationary. As can be observed from the upper portion of Table 1, the real effective rate between the Rand and the US dollar is stationary outside the range of the two threshold points $1=R7.95 and $1=R9.17, whereas it contains a unit root within these two threshold points. Similar inferences can be drawn for the British Pound Euro, the Chinese Renminbi and the Japanese Yen, respectively, with stationary processes being found outside the real exchange rates of £1=R7.95 and £1=R9.17 for the Pound; €1=R10.65 and €1=R12.15 for the Euro; outside the range of ¥1=R1.12 and ¥1=R1.38 for the Renminbi and a much narrower outer band range of ¥1=R0.80 and ¥1=R0.88 for the Yen.

Therefore in summarizing of the results reported in Tale 1, two independent yet simultaneous evidences of PPP behaviour between South Africa and her main trading currency partners can be observed. Firstly, the partial stationarity found in the computed real exchange rates provides our primary validity of the PPP hypothesis. Secondly, the partial unit root process found between the nominal exchange rate variable and the differences in the price indices presents a second indication of or conformity to the PPP hypothesis. However, with regards to the latter case, the evidence presented is merely preliminary and formal cointegration analysis must be conducted in order to avoid spurious results being associated with any estimated PPP regressions. The paper therefore proceeds to perform formal asymmetric cointegration and threshold error correction analysis between South African nominal exchange rates, on one hand, and the differences in domestic and foreign aggregate prices, on the other hand.

4.3 *CO-INTEGRATION ANALYSIS*

Having established that nominal exchange rates and differences in price levels can be partially rendered as being integrated of order one (i.e. I(1)), the paper implements the asymmetric co-integration model of Enders and Silkos (2001), as discussed in the previous section of the paper. Prior to estimating the threshold co-integration and error correction models, we apply a battery of co-integration and error correction tests to the PPP threshold co-integration regressions between nominal exchange rates and the differences in domestic and foreign aggregate prices. As previously mentioned, we apply four generic cointegration tests to the regressions, namely; (1) tests for the stationarity of the co-integration residuals (2) tests for non-spurious co-integration effects (3) tests for asymmetric co-integration effects; and (4) tests for asymmetric error correction mechanisms. In taking a systematic approach to reporting the results, as presented in Table 2; the upper half of Table 2 presents the hypotheses tests on both the TAR and MTAR models with a zero thresholds whereas the bottom half of Table 2 examines these hypotheses on the TAR and MTAR specifications with consistent threshold estimates.

{INSERT TABLE 2 HERE}

In also undertaking a systematic approach to reporting the results presented in Table 2; we can firstly note that the null hypothesis of stationarity in the co-integration residual (i.e. $H_0^{(1)}$) cannot be rejected for all PPP threshold co-integration residuals. We are able to come to this conclusion since all estimates of the equilibrium threshold error terms (i.e. ρ_1 and ρ_2) satisfy the convergence condition of ρ_1, $\rho_2 < 0$. In testing for our second hypothesis of cointegration effects (i.e. $H_0^{(2)} : \rho_1 = \rho_2 = 0$), we note that for all estimated regressions, the F-statistics exceeds that of the critical values for significant levels of at least 10 percent. This implies that we reject the notion of no cointegration between nominal exchange rates and price differentials for all estimated regressions. In turning to the tests of our third hypothesis of asymmetric cointegration effects (i.e. $H_0^{(3)} : \rho_1 = \rho_2$), we reject the null hypothesis of symmetric effects for 7 model specifications. In particular, we obtain F-statistic of 4.35 for the c-TAR model and a F-statistic of 5.01 for the c-MTAR models for US data which are test-statistics which are above their critical values at a 10 percent significance level. For the UK, we find a significant F-statistic of 4.08 which exceeds it's critical value at 10 percent significance level. For China, we obtain F-statistics of 5.71 for the MTAR model and 6.80 for the c-MTAR model which exceed their critical values at a 5 percent significance level. And for Japan, we find a F-statistic of 2.79 for the c-TAR model and 6.98 for the M-TAR and these figures exceed their critical values at minimum of 10 percent significance level. And even more encouraging, when testing the final null hypothesis of no asymmetric error correction effects (i.e. $H_0^{(4)}: \lambda^+ \xi_{t-1}^+ = \lambda^- \xi_{t-1}^-$), we are able to reject the null hypothesis in favour of asymmetric error correction effects for all 7 regressions which previously rejected the null hypothesis of symmetric cointegration effects. Having this evidence of asymmetric cointegration and threshold error correction effects, we therefore proceed to formally estimate the associated (c)TAR-TEC and (c)MTAR-TEC models for the 7 identified regressions, with the estimation results being reported in Table 3.

{INSERT TABLE 3 HERE}

Based on the long-run regression results reported in upper portions of each of the estimates presented in Table 3, we note long-run cointegration elasticities between price differentials and nominal exchange rates of 0.53 for the US, 0.53 for the UK, 1.1 for China and 0.73 for Japan. These estimates imply that a one percentage change in the price differentials between South Africa and the US as well as between South Africa and the UK, is associated with a 0.53 percentage change in the Rand-Dollar as well as Rand-Pound exchange rates. Furthermore, a one percent change price differentials between South Africa and China would results in a 1.1 percent change in the Rand-Yaun exchange rate whereas a one percent change in price differential between South Africa and Japan will result in a 0.73 percent in the Rand-Yen exchange rate. In turning to the results for the threshold equilibrium results we firstly highlight that for the c-TAR-TEC model describing South Africa-US relations, the estimates of the threshold error correction terms ρ_1 and ρ_2 which are coefficients of negative and positive deviations from the equilibrium, are -0.03 and -0.11, respectively. These results imply that positive deviations from the equilibrium are eradicated faster than negative ones. Similarly the ρ_1 and ρ_2 estimates for the c-MTAR-TEC model for the US produces values of -0.02 and -0.11, respectively, hence also implying that positive deviations are eradicated quicker than negative ones. For South Africa-UK relations, the coefficients of ρ_1 and ρ_2 are -0.02 and -0.10, respectively which also means that positive deviations are erased quicker than negative ones. Also for South-Chain relations, the MTAR-TEC model depicts that positive deviations are eradicated quicker than negative ones as can be seen from the estimates of -0.01 and -0.07 for ρ_1 and ρ_2, respectively. Moreover, the c-MTAR-TEC model estimates of ρ_1 and ρ_2 for South-Africa-China relations are -0.01 and -0.09, respectively, hence drawing similar implications of positive deviations being eradicated quicker in comparison to negative ones. For South Africa-Japan relations we find estimates of -0.02 and -0.08 for the MTAR-TEC model and -0.01 and -0.09 for the c-MTAR-TEC model. Once again, both models imply that positive deviations from the steady-state are eradicated quicker than negative ones.

On the other hand, the estimated error correction coefficients ξ_{t-1}^{-} and ξ_{t-1}^{+} reported in lower portions of each of the estimates presented in Table 3, respectively measure the speed of adjustment for negative and positive deviations from the long-run PPP. Furthermore, deviations from the equilibrium level can only be deemed to be self-correcting if at least one the error correction terms in the error correction models is significantly negative. In particular, the negative estimate of the error correction term reveals the speed adjustment at which shocks to either nominal exchange rates or differences in aggregate prices will result in reversion back to equilibrium. Generally, our results indicate that for all estimated regression equations, the only negative and significant error correction terms are found when deviations from the equilibrium are positive with the nominal exchange rate being the driving force in the error correction system. At this juncture it should be noted that these results are in coherence with those presented by Enders and Chumrusphonlert (2004) who, for Asian-pacific economies, find evidence of significant equilibrium reversion behaviour only when the error correction mechanism is being determined by the nominal exchange rates and the deviations are positive. However, in elaborating on the results presented in Table 3, we find for the South African-US case that positive nominal exchange rate shocks converge back to long-run equilibrium at the rate of 11 percent when the shocks are abrupt and at a slightly lower rate of 10 percent when shocks are smooth. The South African-UK case is a particularly interesting case in which we establish relative high equilibrium reversion rates of 90 percent when a positive nominal exchange rate shock is induced in the system. In the case of South African-Chinese PPP relations, mean reversion towards equilibrium is at 6 percent when nominal exchange rate shocks are abrupt and at 8 percent when disequilibrium is smooth whereas for the South African-Japanese case, mean reversion for abrupt shocks is self-correcting at 8 percent and 9 percent for smooth shocks.

Given evidence of threshold cointegration and error correction mechanisms between the exchange rate and differences in price levels, it would be useful to enquire as to whether nominal exchange rates are the endogenous or exogenous variables within the estimated asymmetric PPP relationships. To this end we run granger causality tests on time series variables and report the results in Table 4. The causality tests reveal that, for all estimated equations; nominal exchange rates (i.e. ϵ_t) are deemed to granger cause aggregate price levels (i.e. π_t^*). Although this results contradicts conventional economic wisdom which speculates on causality running from price differentials to exchange rates, notably our results are in coherence with those obtained in Einzig (1935), Hafer (1989), Kholdy and Sohrabian (1990), Menon (1995) and Schnabl and Baur (2002). For the case of South Africa, causality running from exchange rates to price differentials are plausible for a number of reasons. Firstly, the empirical data covers a time period characterized by the disintegration of the Bretton Woods system which led to exchange rate reforms in the South African economy. The structural shifts in exchange rate policies, the global recession periods of 1973 and 1979 as well as the increases in gold price of 1979, all account for the increased price differential experienced between South Africa and her trading partners and not the other way around.

Secondly, political unrest and the resulting sanctions placed on South Africa of 1985 to 1994 forced monetary authorities to use more direct controls in managing the exchange rates. This, in turn, would ensure that exchange rates movements would be responsible for price differentials and not vice versa. Thirdly, following the democratic elections of 1994, the progressive relaxation of exchange controls resulted in improved capital inflows and ultimately improved price differential between South Africa and her trading currency partners would account for causality running from exchange rates to price differentials. Fourthly, the spill over effects from the Asian financial crisis of 1998 caused a weakening of the Rand against her major trading partners which then contributed to increasing inflation prices compared to trading partners thus insinuating that weaker exchange rates led to worsening pricing differentials between South Africa and her trading partners. Fifthly, the infamous global financial crisis of 2008 resulted in upward pressure being placed on the South African Rand as investors sold Rand-dominated assets which contributed to increased inflation rates during this period. And lastly, the South African Rand lost 26 percent of its value in 2015, after the Chinese Central Bank devalued its currency in attempts to boost its export competitiveness. Furthermore, the reshuffling of the Finance ministry in 2015 resulted in undermined market confidence which further contributed to South Africa's currency woes. This, in turn, has resulted in a rise in a weakening fo the South African Rand which then led to an increase in domestic inflation trends in comparison to inflation experienced in South Africa's main trading partners. All-in-all, the aforementioned developments can hold as a suitable explanation in providing a relevant explanation for the causality results obtained in our empirical analysis.

{INSERT TABLE 4 HERE}

5 CONCLUSIONS

In view of a lack of evidence analysing possible asymmetric behaviour in the PPP behaviour between South Africa and her main currency trading partners, namely the US, the UK, the Euro area, China and Japan; our study sought to fill this hiatus in the academic paradigm using a two-stage empirical process conducted on monthly data collected between 1998 and 2013. In the first stage of our empirical analysis, we examine the integration properties of the real exchange rate as computed as the logarithmic transformation of the nominal exchange rates adjusted for price differentials between the South Africa and her trading currency partners. As a point of departure from the common norm of linear unit root test as standardized in the empirical literature; this study opted to apply the nonlinear unit root tests of Kapetanois and Shin (2006) to the empirical data. Empirical evidence showed significant PPP behaviour between South Africa and all her main trading partners, and yet the significance of such PPP behaviour is nonlinear, that is, it only exists outside a specified range of real exchange rates.

In the second stage of our empirical analysis, formal TAR-TEC and MTAR-TEC models were introduced as a means of determining the extent to which nominal exchange rates and the differences in the domestic and foreign aggregate price levels are asymmetrically co-integrated. Our empirical results confirm significant asymmetric cointegration evidence for all South Africa's currency trading partners with the sole exception of the Euro area. In particular, we find that positive deviations from the long-run equilibrium are easier to eradicate in comparison to negative ones for all country cases. This implies that positive developments in exchange rates and price differentials between South Africa and her main currency trading partners are absorbed quicker than negative shocks. In other words, the effects of negative developments in exchange rates or price differentials are likely to last longer than positive ones. Take for instance, during periods following both the Asian financial crisis of 1998 and the global financial crisis of 2008, the weakening of the Rand against the currencies of her trading partners failed for a few years subsequent to the negative shocks to the exchange rates. Conversely, in periods where the exchange rate has significantly appreciated, as is the case in 2005 when South Africa Rand appreciated significant due to a sharp increase in net cash flows, then such positive developments are eradicated quicker than is the case for negative shocks. Our empirical analysis was then supplemented with granger causality tests. Contrary to conventional belief, the granger causality tests revealed that nominal exchange rates are exogenous whereas aggregate prices are endogenous, that is, causality was rendered to solely run from nominal exchange rates to aggregate prices.

In conclusion, our study confirms the importance of the PPP hypothesis for monetary policy conduct in South Africa by placing emphasis on the stability of exchange rates, for not only controlling aggregate price levels, but in also improving the competitive behaviour of domestic prices in international markets. In particular, the empirical analysis reveals that stability in exchange rates can be achieved through stability in aggregate price levels and yet price stability between South Africa and her trading partners will not affect the exchange rate. This result is of particular importance taking into consideration the increasing participation of South African Reserve Bank’s (SARB) involvement in building up foreign exchange reserves as this involves purchasing foreign exchange from financial markets. In terms of policy implications, our results ultimately depict that an exchange rate targeting framework may prove to be a useful avenue for future macroeconomic stabilization policies available to the SARB. Therefore our empirical evidence supplements those presented by Bonga-Bonga and

Kabundi (2010); Phiri (2012) and Gupta (2013) in advocating for the use of a flexible exchange rate targeting frameworks as a more feasible monetary policy alternative as opposed to the current inflation-targeting regime which is under heavy criticism for being a rather rigid monetary policy framework.

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Table 1: Kapetanois and Shin (2006) Unit Test Results

	test statistics									threshold estimat	
	$\mathcal{W}_{sup}$			$\mathcal{W}_{ave}$			$\mathcal{W}_{exp}$				
	none	*intercept*	*trend*	*none*	*intercept*	*trend*	*none*	*intercept*	*trend*	γ_1	γ_2
$\tau_t/us\$$	17.07***	12.87***	6.39	4.25	5.44	4.98	128.60**	63.74*	14.64	794.9	917.2
$\tau_t/uk£$	7.63**	8.94*	16.73***	3.56	4.11	5.47	11.87*	13.91	386.98*	794.9	917.2
$\tau_t/euro€$	12.47***	12.81***	13.22**	6.40*	8.89**	8.93	77.93**	191.50**	208.81*	1065	1215
$\tau_t/china¥$	15.68***	16.18***	21.63***	6.24*	6.74	7.06	89.74**	93.31**	2364.29**	111.9	137.6
$\tau_t/japan¥$	7.13*	18.88***	25.26***	2.63	4.33	2.50	7.73	278.68**	3259.77***	8.00	8.80
$\epsilon_t/us\$$	17.08***	12.88***	6.39	4.26	5.44	4.98	128.94**	63.97*	14.65	795.1	917.5
$\epsilon_t/uk£$	14.64***	8.95*	16.74***	4.48	4.11	5.47	59.48**	13.21	388.47*	795.1	917.5
$\epsilon_t/euro€$	12.47***	12.82***	13.24**	6.40*	8.90*	8.94	36.34**	192.80**	210.06*	1065	1215
$\epsilon_t/china¥$	15.57***	16.06***	21.54***	6.24*	6.73	7.05	88.28**	91.08*	2322.80**	112.2	128.7
$\epsilon_t/japan¥$	8.19**	21.48***	24.56***	3.32	5.01	2.48	13.08*	674.18***	2323.79**	8.44	9.80
π_t^*/us	10.45**	8.73*	12.14*	4.15	7.94*	5.97	22.53**	54.27*	68.51	0.1958	0.2339
π_t^*/uk	14.63***	10.85**	14.11**	4.47	7.07	8.30	59.04**	82.93*	152.57	0.1869	0.2008
$\pi_t^*/euro$	10.24**	15.16***	14.22**	6.18*	3.88	8.53	36.38**	332.01**	482.83***	0.2094	0.3031
$\pi_t^*/china$	7.24*	12.02**	13.89**	3.23	6.91	11.29*	9.47*	93.56**	450.68**	0.3220	0.3455
$\pi_t^*/japan$	7.82**	10.11*	25.26***	1.70	5.52	2.50	7.08	48.29*	3259.77**	0.2875	0.3484
10%	6.01	7.29	10.35	6.01	7.29	10.35	7.49	38.28	176.80		
critical 5%	7.49	9.04	12.16	7.49	9.04	12.16	20.18	91.83	437.03		
values 1%	10.49	12.64	16.28	10.49	12.64	16.28	237.46	555.57	3428.92		

Significance Level Codes are as follows."***", "**' and '*' denote the 1%, 5% and 10% significance levels respectively. The variables $\tau_t/us\$$ $\tau_t/uk£$, $\tau_t/euro€$ $\tau_t/china¥\tau_t/japan¥$ represent the computed real effective exchange rate between the US dollar, the British Pound, the Euro, the Chinese Renminbi and the Japanese Yen, respectively. The variables $\varepsilon_t/us\$$, $\varepsilon_t/uk£$, $\varepsilon_t/euro€$, $\varepsilon_t/china¥\tau_t/japan¥$ represent the nominal exchange rate between the Rand and the US dollar, the British Pound, the Euro, the Chinese Renminbi and the Japanese Yen, respectively. The variables π_t/us π_t/uk, $\pi_t/euro$, $\pi_t/china, \pi_t/japan$ represent the nominal exchange rate between the Rand and the US dollar, the British Pound, the Euro, the Chinese Renminbi and the Japanese Yen, respectively. The test statistics $\mathcal{W}_{sup}\mathcal{W}_{ave}\mathcal{W}_{exp}$are the supremum, average and exponential average variations of the Wald statistic.

Table 2: Co-integration and error correction tests for TAR and c-TAR models

	dependent variable Y	*independent variable* X	*TAR/c − TAR* $H_0^{(1)}$	$H_0^{(2)}$	$H_0^{(3)}$	$H_0^{(4)}$	*MTAR/c − MTAR* $H_0^{(1)}$	$H_0^{(2)}$	$H_0^{(3)}$	$H_0^{(4)}$
zero threshold	$\epsilon_t/us\$$	π_t^*/us	*reject*	4.30	1.71	2.23	*reject*	4.62	2.31	3.11
			null	$(0.01)^{***}$	(0.19)	$(0.03)^{*}$	*null*	$(0.01)^{***}$	(0.13)	$(0.08)^{*}$
	$\epsilon_t/uk£$	π_t^*/uk	*reject*	3.72	0.41	2.70	*reject*	4.36	1.65	1.76
			null	$(0.03)^{*}$	(0.52)	$(0.10)^{*}$	*null*	$(0.01)^{***}$	(0.20)	(0.19)
	$\epsilon_t/euro€$	$\pi_t^*/euro$	*reject*	4.02	0.11	4.61	*reject*	3.98	0.03	0.01
			null	$(0.02)^{**}$	(0.74)	$(0.03)^{*}$	*null*	$(0.02)^{**}$	(0.85)	(0.92)
	$\epsilon_t/china¥$	$\pi_t^*/china$	*reject*	2.36	0.11	0.03	*reject*	5.23	5.71	4.42
			null	$(0.10)^{*}$	(0.74)	(0.86)	*null*	$(0.01)^{***}$	$(0.02)^{**}$	$(0.04)^{**}$
	$\epsilon_t/japan¥$	$\pi_t^*/japan$	*reject*	4.06	0.29	1.03	*reject*	5.36	2.79	2.16
			null	$(0.02)^{**}$	(0.59)	(0.31)	*null*	$(0.01)^{***}$	$(0.10)^{*}$	$(0.14)^{*}$
consistent threshold estimate	$\epsilon_t/us\$$	π_t^*/us	*reject*	5.67	4.35	4.32	*reject*	6.02	5.01	4.51
			null	$(0.00)^{***}$	$(0.04)^{*}$	$(0.04)^{*}$	*null*	$(0.00)^{***}$	$(0.03)^{*}$	$(0.04)^{*}$
	$\epsilon_t/uk£$	π_t^*/uk	*reject*	4.13	1.21	1.25	*reject*	5.62	4.08	4.55
			null	$(0.02)^{**}$	(0.27)	(0.27)	*null*	$(0.00)^{***}$	$(0.05)^{*}$	$(0.03)^{*}$
	$\epsilon_t/euro€$	$\pi_t^*/euro$	*reject*	4.16	0.39	3.67	*reject*	5.09	2.16	2.60
			null	$(0.02)^{**}$	(0.53)	$(0.06)^{*}$	*null*	$(0.01)^{***}$	(0.14)	$(0.11)^{*}$
	$\epsilon_t/china¥$	$\pi_t^*/china$	*reject*	2.48	0.34	0.29	*reject*	5.79	6.80	5.62
			null	$(0.09)^{*}$	(0.56)	(0.59)	*null*	$(0.00)^{***}$	$(0.01)^{***}$	$(0.02)^{**}$
	$\epsilon_t/japan¥$	$\pi_t^*/japan$	*reject*	4.43	1.01	2.02	*reject*	7.54	6.98	6.84
			null	$(0.01)^{***}$	(0.32)	(0.16)	*null*	$(0.00)^{***}$	$(0.01)^{***}$	$(0.01)^{***}$

Significance Level Codes:"***", "**' and '*' denote the 1%, 5% and 10% significance levels respectively. The p-values are reported in ().

$H_0^{(1)}, H_0^{(2)}, H_0^{(3)}$ *and* $H_0^{(4)}$ are hypotheses testing for stationarity of the equilibrium error term; normal co-integration effects; asymmetric co-integration effects; and asymmetric error correction effects, respectively.

Table 3: (c)TAR-TEC and (c)MTAR-TEC Model estimates

For US

$c-TAR-TEC$

$$\epsilon_t = \underset{(0.00)^{***}}{6.58} + \underset{(0.00)^{***}}{0.53}\,\pi_t^* - \underset{(0.17)}{0.03}\xi_{t-1}I.(\xi_{t-1} < -0.13) - \underset{(0.00)^{***}}{0.11}\xi_{t-1}I.(\xi_{t-1} \geq -0.13) + \underset{(0.00)^{***}}{0.33}\,\Delta\xi_{t-i}$$

$$\Delta\pi_t^* = \underset{(0.01)^{***}}{0.02} + \begin{cases} \underset{(0.08)^{*}}{0.16}\,\Delta\pi_{t-1}^* + \underset{(0.00)^{***}}{0.05}\,\Delta\epsilon_{t-1} + \underset{(0.00)^{***}}{0.01}\,\xi_{t-1}I.(\xi_{t-1} \geq -0.13) \\ \underset{(0.21)}{0.26}\,\Delta\pi_{t-1}^* + \underset{(0.27)}{0.02}\,\Delta\epsilon_{t-1} + \underset{(0.90)}{0.00}\,\xi_{t-1}I.(\xi_{t-1} < -0.13) \end{cases}$$

$$\Delta\epsilon_t = \underset{(0.40)}{0.01} + \begin{cases} \underset{(0.78)}{0.21}\,\Delta\pi_{t-1}^* + \underset{(0.02)^{**}}{0.26}\,\Delta\epsilon_{t-1} - \underset{(0.41)}{0.02}\,\xi_{t-1}I.(\xi_{t-1} \geq -0.13) \\ \underset{(0.78)}{0.49}\,\Delta\pi_{t-1}^* + \underset{(0.00)^{***}}{0.52}\,\Delta\epsilon_{t-1} - \underset{(0.01)^{***}}{0.11}\,\xi_{t-1}I.(\xi_{t-1} < -0.13) \end{cases}$$

$c-MTAR-TEC$

$$\epsilon_t = \underset{(0.00)^{***}}{6.58} + \underset{(0.00)^{***}}{0.53}\,\pi_t^* - \underset{(0.22)}{0.02}\xi_{t-1}I.(\Delta\xi_{t-1} < -0.02) - \underset{(0.00)^{***}}{0.11}\xi_{t-1}I.(\Delta\xi_{t-1} \geq -0.02) + \underset{(0.00)^{***}}{0.32}\,\Delta\xi_{t-i}$$

$$\Delta\pi_t^* = \underset{(0.00)^{*}}{0.00} + \begin{cases} \underset{(0.09)^{*}}{0.15}\,\Delta\pi_{t-1}^* + \underset{(0.00)^{***}}{0.05}\,\Delta\epsilon_{t-1} + \underset{(0.03)^{*}}{0.01}\,\xi_{t-1}I.(\Delta\xi_{t-1} \geq -0.02) \\ \underset{(0.10)^{*}}{0.30}\,\Delta\pi_{t-1}^* + \underset{(0.29)}{0.02}\,\Delta\epsilon_{t-1} + \underset{(0.11)^{*}}{0.01}\,\xi_{t-1}I.(\Delta\xi_{t-1} < -0.02) \end{cases}$$

$$\Delta\epsilon_t = \underset{(0.17)}{0.01} + \begin{cases} \underset{(0.78)}{-0.22}\,\Delta\pi_{t-1}^* + \underset{(0.02)^{**}}{0.25}\,\Delta\epsilon_{t-1} - \underset{(0.37)}{0.02}\,\xi_{t-1}I.(\Delta\xi_{t-1} \geq -0.02) \\ \underset{(0.63)}{0.82}\,\Delta\pi_{t-1}^* + \underset{(0.00)^{***}}{0.50}\,\Delta\epsilon_{t-1} - \underset{(0.00)^{***}}{0.10}\,\xi_{t-1}I.(\Delta\xi_{t-1} < -0.02) \end{cases}$$

For UK

$c-MTAR-TEC$

$$\epsilon_t = \underset{(0.00)^{***}}{7.10} + \underset{(0.00)^{***}}{0.53}\,\pi_t^* - \underset{(0.56)}{0.02}\xi_{t-1}I.(\Delta\xi_{t-1} \geq -0.01) - \underset{(0.00)^{***}}{0.10}\xi_{t-1}I.(\Delta\xi_{t-1} < -0.01) + \underset{(0.01)^{***}}{0.18}\,\Delta\xi_{t-i}$$

$$\Delta\pi_t^* = \underset{(0.01)^{***}}{0.00} + \begin{cases} -\underset{(0.02)^{**}}{0.21}\,\Delta\pi_{t-1}^* + \underset{(0.19)}{0.03}\,\Delta\epsilon_{t-1} + \underset{(0.01)^{***}}{0.01}\,\xi_{t-1}I.(\Delta\xi_{t-1} \geq -0.01) \\ \underset{(0.39)}{0.23}\,\Delta\pi_{t-1}^* + \underset{(0.12)^{*}}{0.04}\,\Delta\epsilon_{t-1} + \underset{(0.01)^{***}}{0.01}\,\xi_{t-1}I.(\Delta\xi_{t-1} < -0.01) \end{cases}$$

$$\Delta\epsilon_t = \underset{(0.15)}{0.01} + \begin{cases} -\underset{(0.83)}{0.12}\,\Delta\pi_{t-1}^* + \underset{(0.52)}{0.07}\,\Delta\epsilon_{t-1} - \underset{(0.98)}{0.00}\,\xi_{t-1}I.(\Delta\xi_{t-1} \geq -0.01) \\ \underset{(0.95)}{-0.10}\,\Delta\pi_{t-1}^* + \underset{(0.02)^{**}}{0.40}\,\Delta\epsilon_{t-1} - \underset{(0.01)^{***}}{0.90}\,\xi_{t-1}I.(\Delta\xi_{t-1} < -0.01) \end{cases}$$

For CHINA

$MTAR-TEC$

$$\epsilon_t = \underset{(0.00)^{***}}{4.41} + \underset{(0.00)^{***}}{1.11}\,\pi_t^* - \underset{(0.71)}{0.01}\xi_{t-1}I.(\Delta\xi_{t-1} \geq 0) - \underset{(0.00)^{***}}{0.07}\xi_{t-1}I.(\Delta\xi_{t-1} < 0) + \underset{(0.01)^{***}}{0.31}\,\Delta\xi_{t-i}$$

$$\Delta\pi_t^* = \underset{(0.56)}{0.00} + \begin{cases} \underset{(0.00)^{***}}{0.37}\,\Delta\pi_{t-1}^* - \underset{(0.08)^{*}}{0.04}\,\Delta\epsilon_{t-1} + \underset{(0.70)}{0.00}\,\xi_{t-1}I.(\Delta\xi_{t-1} \geq 0) \\ \underset{(0.65)}{-0.07}\,\Delta\pi_{t-1}^* + \underset{(0.02)^{**}}{0.02}\,\Delta\epsilon_{t-1} + \underset{(0.00)^{***}}{0.02}\,\xi_{t-1}I.(\Delta\xi_{t-1} < 0) \end{cases}$$

$$\Delta\epsilon_t = \underset{(0.04)^{**}}{0.01} + \begin{cases} \underset{(0.65)}{-0.22}\,\Delta\pi_{t-1}^* + \underset{(0.06)^{*}}{0.21}\,\Delta\epsilon_{t-1} + \underset{(0.55)}{0.02}\,\xi_{t-1}I.(\Delta\xi_{t-1} \geq 0) \\ \underset{(0.48)}{0.53}\,\Delta\pi_{t-1}^* + \underset{(0.00)^{***}}{0.60}\,\Delta\epsilon_{t-1} - \underset{(0.01)^{***}}{0.06}\,\xi_{t-1}I.(\Delta\xi_{t-1} < 0) \end{cases}$$

$$\epsilon_t = \underset{(0.00)^{***}}{4.41} + \underset{(0.00)^{***}}{1.11}\pi_t^* - \underset{(0.93)}{0.01}\xi_{t-1}I.(\Delta\xi_{t-1} \geq -0.01) - \underset{(0.00)^{***}}{0.09}\xi_{t-1}I.(\Delta\xi_{t-1} < -0.01) + \underset{(0.01)^{***}}{0.32}\Delta\xi_{t-i}$$

$$\Delta\pi_t^* = \underset{(0.53)}{0.00} + \begin{cases} \underset{(0.00)^{***}}{0.38}\Delta\pi_{t-1}^* + \underset{(0.12)}{0.04}\Delta\epsilon_{t-1} + \underset{(0.25)}{0.01}\xi_{t-1}I.(\Delta\xi_{t-1} \geq -0.01) \\ \underset{(0.58)}{-0.09}\Delta\pi_{t-1}^* + \underset{(0.59)}{0.02}\Delta\epsilon_{t-1} + \underset{(0.01)^{***}}{0.02}\xi_{t-1}I.(\Delta\xi_{t-1} < -0.01) \end{cases}$$

$c - MTAR - TEC$

$$\Delta\epsilon_t = \underset{(0.05)^{**}}{0.01} + \begin{cases} \underset{(0.67)}{-0.21}\Delta\pi_{t-1}^* + \underset{(0.04)^{*}}{0.22}\Delta\epsilon_{t-1} + \underset{(0.80)}{0.01}\xi_{t-1}I.(\Delta\xi_{t-1} \geq -0.01) \\ \underset{(0.42)}{0.60}\Delta\pi_{t-1}^* + \underset{(0.00)^{***}}{0.61}\Delta\epsilon_{t-1} - \underset{(0.01)^{***}}{0.08}\xi_{t-1}I.(\Delta\xi_{t-1} < -0.01) \end{cases}$$

For JAPAN

$MTAR - TEC$

$$\epsilon_t = \underset{(0.00)^{***}}{1.91} + \underset{(0.00)^{***}}{0.73}\pi_t^* - \underset{(0.53)}{0.02}\xi_{t-1}I.(\Delta\xi_{t-1} \geq 0) - \underset{(0.00)^{***}}{0.08}\xi_{t-1}I.(\Delta\xi_{t-1} < 0) + \underset{(0.00)^{***}}{0.28}\Delta\xi_{t-i}$$

$$\Delta\pi_t^* = \underset{(0.00)^{***}}{0.00} + \begin{cases} \underset{(0.49)}{0.06}\Delta\pi_{t-1}^* + \underset{(0.02)^{**}}{0.03}\Delta\epsilon_{t-1} - \underset{(0.57)}{0.02}\xi_{t-1}I.(\Delta\xi_{t-1} \geq 0) \\ \underset{(0.00)^{***}}{1.09}\Delta\pi_{t-1}^* + \underset{(0.81)}{0.01}\Delta\epsilon_{t-1} + \underset{(0.03)^{**}}{0.01}\xi_{t-1}I.(\Delta\xi_{t-1} < 0) \end{cases}$$

$$\Delta\epsilon_t = \underset{(0.60)}{0.00} + \begin{cases} \underset{(0.72)}{-0.26}\Delta\pi_{t-1}^* + \underset{(0.01)^{**}}{0.30}\Delta\epsilon_{t-1} - \underset{(0.57)}{0.02}\xi_{t-1}I.(\Delta\xi_{t-1} \geq 0) \\ \underset{(0.69)}{-1.26}\Delta\pi_{t-1}^* + \underset{(0.10)^{*}}{0.28}\Delta\epsilon_{t-1} - \underset{(0.01)^{***}}{0.08}\xi_{t-1}I.(\Delta\xi_{t-1} < 0) \end{cases}$$

$c - MTAR - TEC$

$$\epsilon_t = \underset{(0.00)^{***}}{1.91} + \underset{(0.00)^{***}}{0.73}\pi_t^* - \underset{(0.66)}{0.01}\xi_{t-1}I.(\Delta\xi_{t-1} \geq 0.02) - \underset{(0.00)^{***}}{0.09}\xi_{t-1}I.(\Delta\xi_{t-1} < 0.02) + \underset{(0.00)^{***}}{0.27}\Delta\xi_{t-i}$$

$$\Delta\pi_t^* = \underset{(0.00)^{***}}{0.00} + \begin{cases} \underset{(0.29)}{0.09}\Delta\pi_{t-1}^* + \underset{(0.03)^{**}}{0.03}\Delta\epsilon_{t-1} + \underset{(0.10)^{*}}{0.01}\xi_{t-1}I.(\Delta\xi_{t-1} \geq 0.02) \\ \underset{(0.01)^{***}}{1.03}\Delta\pi_{t-1}^* + \underset{(0.91)}{0.00}\Delta\epsilon_{t-1} + \underset{(0.03)^{*}}{0.00}\xi_{t-1}I.(\Delta\xi_{t-1} < 0.02) \end{cases}$$

$$\Delta\epsilon_t = \underset{(0.69)}{0.00} + \begin{cases} \underset{(0.85)}{-0.13}\Delta\pi_{t-1}^* + \underset{(0.01)^{***}}{0.28}\Delta\epsilon_{t-1} + \underset{(0.55)}{0.20}\xi_{t-1}I.(\Delta\xi_{t-1} \geq 0.02) \\ \underset{(0.64)}{-1.41}\Delta\pi_{t-1}^* + \underset{(0.10)^{*}}{0.28}\Delta\epsilon_{t-1} - \underset{(0.00)^{***}}{0.09}\xi_{t-1}I.(\Delta\xi_{t-1} < 0.02) \end{cases}$$

Significance Level Codes:"***", "**' and '*' denote the 1%, 5% and 10% significance levels respectively. The associated p-values are reported in parentheses ().

Table 4: Granger Causality tests

$country$	$model\ type$	Y	X	$H_0^{(5)}$	$H_0^{(6)}$	$decision$
RSA/US	$c-TAR$ $-TEC$	ϵ_t	π_t^*	0.06 (0.94)	14.13 $(0.00)^{***}$	$\epsilon_t\ to\ \pi_t^*$
	$c-MTAR$ $-TEC$	ϵ_t	π_t^*	0.12 (0.88)	12.69 $(0.00)^{***}$	$\epsilon_t\ to\ \pi_t^*$
RSA/UK	$c-MTAR$ $-TEC$	ϵ_t	π_t^*	0.04 (0.96)	4.51 $(0.01)^{**}$	$\epsilon_t\ to\ \pi_t^*$
$RSA/CHINA$	$c-MTAR$ $-TEC$	ϵ_t	π_t^*	0.27 (0.76)	14.39 $(0.00)^{***}$	$\epsilon_t\ to\ \pi_t^*$
	$MTAR$ $-TEC$	ϵ_t	π_t^*	0.33 (0.72)	15.18 $(0.00)^{***}$	$\epsilon_t\ to\ \pi_t^*$
$RSA/JAPAN$	$c-MTAR$ $-TEC$	ϵ_t	π_t^*	0.22 (0.80)	8.48 $(0.00)^{***}$	$\epsilon_t\ to\ \pi_t^*$
	$MTAR$ $-TEC$	ϵ_t	π_t^*	0.17 (0.84)	7.90 $(0.00)^{***}$	$\epsilon_t\ to\ \pi_t^*$

Significance Level Codes:"***", "**' and '*' denote the 1%, 5% and 10% significance levels respectively. P-values are reported in parentheses (). $H_0^{(5)}$ is the null hypothesis that the price differentials do not lead to nominal exchange rate whereas $H_0^{(6)}$ is the null hypothesis that the nominal exchange rates do not lead to price differentials.