Purchasing power parity (PPP) between South Africa and her main currency exchange partners: Evidence from asymmetric unit root tests and threshold co-integration analysis

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PURCHASING POWER PARITY (PPP) BETWEEN SOUTH ARFICA AND HER MAIN CURRENCY EXCHANGE PARTNERS: EVIDENCE FROM ASYMMETRIC UNIT ROOT TESTS AND THRESHOLD CO-INTEGRATION ANALYSIS

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ABSTRACT:

Purpose: This purpose of this study is to examine the asymmetric adjustment effects for the purchasing power parity (PPP) for South Africa against her main currency trading partners; namely, the US, the UK, the Euro area, China and Japan.

Design/Methodology/Approach: This study presents a two-fold empirical approach by using nominal exchange rate and aggregate price level data collected monthly for the periods 1971-2013. As a first step, the paper tests for nonlinear integration properties on the real exchange rate as computed as the nominal exchange rate adjusted for price differentials between the domestic and foreign price levels. The paper then proceeds to investigate asymmetric cointegration and error correction effects between nominal exchange rates and aggregate price differentials; and further supplements the empirical analysis by investigating granger causal effects between the variables.

Findings: While the study is able to validate significant asymmetric PPP effects between South Africa and all her main currency exchange partners through the application of asymmetric unit root tests; the evidence presented when examining these PPP effects through the use of threshold cointegration and error correction analysis exempts the relationship explored between South African and the Euro area. Furthermore, the causal effects are found to run uni-directional from exchange rates to aggregate price differentials for all significant asymmetric cointegration relations.

Originality/value: This study makes a novel contribution to literature by confirm significant asymmetric PPP effects between South Africa and her main currency exchange partners from both a unit root and a co-integration perspective.

Keywords: Purchasing power parity (PPP), Threshold co-integration, Threshold unit root tests, South Africa.

JEL Classification Code: B22, 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, such that foreign currencies should possess the same purchasing power and, consequentially, 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). Historically, the intellectual origins of the PPP theory has been unanimously attributed to the early pioneering work of Cassel (1916) although some commentators highlight that the theory may have emerged as early as the 15th century when Spanish theologian Domingo Banez (1527-1604) and other Salamanca scholars argued that currencies exchange at a parity allowing for the same purchase of the same basket between two economies. Regardless of the true origins of the theory, the empirical validity of the PPP hypothesis, nevertheless, bears important financial implications from an empirical, a theoretical as well as policy perspectives.

From an empirical standpoint, 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) and based upon the existing literature, there are three key motivations as to why the empirical validity of the PPP hypothesis plays an important role in the academic paradigm. Firstly, since the theory of PPP is viewed as a theory of exchange rate determination, then stationarity of the real effective exchange rate is quite essential; since a real exchange rate which is characterized by a unit root implies that innovations to the real exchange rate are highly persistent and the time series can fluctuate without bound (Cashin et. al., 2004). Secondly, given that much of the open or external macroeconomy policy is based on the PPP hypothesis, failure to uphold a stationary real exchange rate series will offer a reason to put into question open economy macroeconomic theory (Narayan, 2005). Thirdly, stationarity of the real exchange rate has implications for practical purposes since estimates of the PPP are frequently used in determining the degree of misalignment of the nominal exchange rate; the appropriate policy response to detected misalignments in the exchange rate; the setting of exchange parities, and international comparison of national income levels (Taylor and Taylor, 2004). In this regard, the PPP hypothesis is viewed as a suitable prediction model which allows policymakers to discern as to whether or not their exchange rate is overvalued or undervalued (Narayan et. al., 2009).

The empirical confirmation of the PPP hypothesis, in turn, bears a number of important implications towards practical policy analysis. For instance, Rogoff (1996) notes that the degree of persistence in the real exchange rate can be used to infer the principal impulses driving exchange rate movements. Furthermore, Bararumshah et. al. (2010) argue that since the real exchange rate is frequently considered as an important measure of international competiveness, more particularly for emerging economies where exports are a principal source of economic growth, monetary authorities in developing economies are typically concerned about large and persistent deviations from the PPP since exchange rates are likely to affect net exports, as well as the cost of servicing foreign-currency-denominated-debt. In other words, the validity of PPP is of particular importance to policymakers in developing economies since the PPP provides an important basis for financial stability and structural adjustment policies which are designed to improve the external competitiveness and is consequentially used as a measure of economic integration (Liu and Su, 2011). The PPP is also viewed as a useful policy tool in determining the design of monetary policy and assessing whether a flexible exchange rate system is successful in insulating the domestic economy from foreign shocks (Frankel, 1981). Thus when the PPP hypothesis does not hold, the use of the monetary approach in determining the exchange rate level is invalidated as this approach requires that the PPP holds true (Bozoklu and Kutlu, 2012).

Even though there exists an almost unanimous agreement that the PPP does not provide a good description of short-run exchange rate movements, no definite evidence has been found as to whether PPP holds in the long run (Oh, 1996). For instance, Brissimis et. al. (2005) argue that the failure of the PPP hypothesis to the hold in the short-run became obvious in the years immediately following the monetary policy shift to floating exchange rates as experienced by a number of Central Banks worldwide, and henceforth, the PPP hypothesis has been accepted as a parity condition linking relative prices and the exchange rate in the long-run. However, even in attempting to model long-run movements in real exchange rates, such attempts by researchers have typically been met with mixed results, more prominently for developing economies as has been demonstrated in a recent publication by Liu et. al. (2011). Therefore, the examination of nonlinear adjustment toward long-run PPP has emerged as an attractive alternative and this empirical view is highly justified on a number of theoretical grounds. Take for instance, Nakagawa (2010) who 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. Other theoretical justifications have been presented by Bozoklu and Kutlu (2012) who put forth claims 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. Furthermore, 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; Su et. al., 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 justified empirical or methodological rationale, preferred one approach over the other but rarely do economists opt to examine or use both approaches simultaneously, let alone from an asymmetric perspective. Our study therefore contributes to the existing literature by applying both asymmetric unit root tests and threshold co-integration analysis to the PPP hypothesis for the South African economy relative to her trading currency partners namely; the United States (US); the United Kingdom (UK); the Euro area; China and Japan between the period of 1998 and 2013. This can be considered a worthwhile contribution to the academic literature 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. Besides, the current literature provides no empirical attempts which test for the causal effects in the PPP relationship for South African data relative to her trading currency partners of which our study conveniently accounts for.

Having provided an introduction and motivation for this current study; the remainder of the paper is outlined 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 of the paper presents an outline of 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 presents the time series data used in the study as well as the empirical analysis as performed on the utilized data. Section five of the paper concludes with policy implications of the results obtained.

2 *PPP AND UNIT ROOT TESTS*

According to Darby (1980), Haikko (1992) and Taylor and Taylor (2004) there are two distinct concepts 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 also be viewed as an extension of the quantity theory of money to an international 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). Empirically, the absolute version of the PPP hypothesis typically assumes the following functional form:

$$\mathcal{P}_t^d = \epsilon_t \mathcal{P}_t^f \qquad (1)$$

Where ϵ_t is the nominal exchange rate, which is commonly defined as the unit price of foreign currency in terms of home currency; and $\mathcal{P}_t^d$ and $\mathcal{P}_t^f$ are the local and foreign price levels, respectively. Despite its simplicity and considerable appeal as theory of equilibrium exchange rates, however, in practice the absolute version of the PPP hypothesis has generally failed on the account of three fundamental reasons. Firstly, the absolute PPP theory seemingly holds only when the purchasing power of a unit for currency is exactly equal in the domestic economy and the foreign economy, once it is converted into foreign currency at the market exchange rate (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 these irregularities (Shirley, 2013).

Due to the preceding 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, such that the relative PPP unconditional holds when the absolute PPP holds, and yet 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 (Kargbo, 2006) and consequentially, researchers commonly opt to use a logarithmic version of the PPP hypothesis as specified below:

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

Due to transaction costs and other impediments caused by trade restrictions, research academics commonly relax the homogeneity restrictions (i.e. $\alpha = 0$) as well as the symmetry and proportionality restrictions (i.e. $\beta_1 = -\beta_2$ and $\beta_1 = -\beta_2 = 1,$ respectively). 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:

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

Bahmani-Oskooee and Gelan (2006) note that under a floating exchange rate system, as adopted by the South African Reserve Bank (SARB), a country's currency could depreciate against one currency and appreciate against another and thus rendering it more feasible to use the real effective exchange rate in examining unit roots in the exchange rate of any given economy. 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. $\frac{\epsilon_t}{\tau_t} = \frac{\mathcal{P}_t^d}{\mathcal{P}_t^f}$) into equation (2) and by re-arranging the terms and further converting the variables into logarithmic form, we can obtain the following PPP regression equation:

$$log\tau_t = log\epsilon_t - \beta log\pi_t^* \tag{4}$$

From the equation (4), the real exchange rate, τ_t, may be, for convenience sake, interpreted as a measure of deviation from PPP equilibrium. In order to validate the PPP hypothesis, one can test whether the real exchange rate is stationary. In testing for stationarity, the real exchange rate can be placed subject to the following generalized autoregressive (Dickey-Fuller-type) regression:

$$\tau_t = \phi\tau_{t-1} + \xi_t \tag{5}$$

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 real exchange rate time series can be modelled as a mean reverting autoregressive process and deviations from the PPP are temporary. Earlier studies that sought to test for stationary of the real exchange rates mostly relied on standard unit root tests such as the augmented Dickey Fuller (ADF) tests and provided little support for the PPP hypothesis (Bahmani-Oskooee and Gelan, 2006). One reason why this may have occurred is that if data generating process is indeed nonlinear, then linear unit root tests will have very low power to reject a false null hypothesis of a unit root and as a means of circumventing this problem, the observed time series variables may require to be tested for unit roots using nonlinear unit root tests. For instance, Taylor (2001) and Bec et. al. (2004) apply nonlinear unit root test to two-regime self exciting threshold autoregressive (SETAR) models for European exchange rates and demonstrate that even though consistent, the standard Dickey-Fuller test lacks power against nonlinear stationary alternatives when the data generating process of exchange rates is indeed nonlinear. Bec et. al. (2004) reach similar conclusions to Taylor (2001) for European exchange rates but opt to apply nonlinear unit root testing procedures under a three-regime SETAR model framework. Kapetanois and Shin (2006) extend on the aforementioned empirical framework by imposing a theoretically congruent condition of the corridor or middle regime of a 3-regime SETAR model as being characterized by an 'inaction band' and thus propose the test procedure for the joint significance of all autoregressive parameters in both inner and outer regimes. In investigating the integration properties of the observed time series, our study follows that of Kapetanois and Shin (2006) by focusing on developing unit root tests based on following nonlinear auxiliary SETAR model:

$$\Delta Y = X(\gamma)\phi + \text{v} \quad (6)$$

Where:

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

The threshold functions for the first and last regimes (third regimes) are given by $\gamma_1 = I.\{\epsilon_i \leq \gamma_1\}$ and $\gamma_2 = I.\{\epsilon_i > \gamma_2\}$, respectively; with γ_1 and γ_2 denoting the first and second threshold estimates, 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: $\mathcal{W}_{\gamma_1,\gamma_2} = \hat{\phi}'[Var(\hat{\phi})]^{-1}\hat{\phi}$ 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 (i.e. $\mathcal{W}_{sup}$), average (i.e. $\mathcal{W}_{ave}$) and exponential average (i.e. $\mathcal{W}_{exp}$) variations of the Wald statistic. 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). Thus 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 which can be expressed as follows:

$$\square_t = log\epsilon_t - log\pi_t^* \qquad (7)$$

From regression (7), 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 $\square_t$~I(0); such that the nominal exchange rates and the differences in price indexes should increase monotonically over time with $\square_t$ 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, $\square_t$, 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/subside 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 the linear co-integration tests fall under an asymmetric adjustment processes and 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. All-in-all, 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 as the follows:

$$\Delta \xi_t = I_t \rho_1 \xi_{t-1} + (1 - I_t)\rho_2 \xi_{t-1} + \sum_{i=1}^{p} \beta_i \, \Delta\xi_{t-i} + \varepsilon_t \quad (8)$$

And thereafter apply the following co-integration tests for (i) stationarity of the equilibrium error term (i.e. $H_0^{(1)} : \rho_1, \rho_2 < 0$) (ii) normal co-integration effects (i.e. $H_0^{(2)} : \rho_1 = \rho_2 = 0$); and (iii) asymmetric co-integration effects (i.e. $H_0^{(3)} : \rho_1 = \rho_2$). 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:

$$I_t = \begin{cases} 1, if\, \xi_{t-1} \geq 0 \\ 0, if\, \xi_{t-1} < 0 \end{cases} \qquad I_t = \begin{cases} 1, if\, \xi_{t-1} \geq q \\ 0, if\, \xi_{t-1} < q \end{cases} \quad (9)$$

The second function 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:

$$M_{.t} = \begin{cases} 1, if \Delta\xi_{t-1} \geq 0 \\ 0, if \Delta\xi_{t-1} < 0 \end{cases} \qquad M_{.t} = \begin{cases} 1, if \Delta\xi_{t-1} \geq q \\ 0, if \Delta\xi_{t-1} < q \end{cases} \tag{10}$$

Since the threshold variable under the c-TAR and c-MTAR models, 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-VEC model can be expressed as:

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

$$\Delta\pi_t^* = \lambda_{21} I_t \xi_{t-1} + \lambda_{22}(1 - I_t)\xi_{t-1} + \sum_{i=1}^{p} \varphi_{i2}\Delta\epsilon_{t-1} + \sum_{i=1}^{p} \psi_{i2}\Delta\pi_{t-1}^* + \mathrm{v}_{t2} \tag{12}$$

Whereas the MTAR-TEC model is specified as:

$$\Delta\epsilon_t = \lambda_{11} M_t \xi_{t-1} + \lambda_{12}(1 - M_t)\xi_{t-1} + \sum_{i=1}^{p} \varphi_{i1} \Delta\epsilon_{t-1} + \sum_{i=1}^{p} \psi_{i1} \Delta\pi_{t-1}^{*} + \mathrm{v}_{t1} \qquad (13)$$

$$\Delta\pi_t^{*} = \lambda_{21} M_t \xi_{t-1} + \lambda_{22}(1 - M_t)\xi_{t-1} + \sum_{i=1}^{p} \varphi_{i2} \Delta\epsilon_{t-1} + \sum_{i=1}^{p} \psi_{i2} \Delta\pi_{t-1}^{*} + \mathrm{v}_{t2} \qquad (14)$$

And the indicator functions as given in regressions (9) and (10) 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 examined 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: $H_0^{(4)}: \lambda^{+} \xi_{t-1}^{+} = \lambda^{-} \xi_{t-1}^{-}$; Whereas, the null hypothesis that the price differentials do not lead to nominal exchange rate is tested as: $H_0^{(5)}: \alpha_k = 0;\ i = 1, \dots, k$ 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: $H_0^{(6)}: \beta_k = 0;\ i = 1, \dots, k$.

4 *EMPIRICAL ANALYSIS*

4.1 *Data Description*

Our data set comprises of a total of 190 monthly observations collected between the periods of January 1998 to October 2013. 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 (is restricted) from the period of 1998 onwards. The data used in our empirical analysis comprises of the time series variables 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 utilized data is transformed into logarithmic form a prior.

A perfunctory observation of the utilized data reveals a number of noteworthy stylized facts which would provide preliminary motivation for the use of asymmetric econometric techniques in analysing the PPP relationship between South Africa and her main currency trading partners. Firstly, we note that our data collection begins during 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 are considered as important perfunctory observations since these shifts in policy conduct may further motivate the need to account for asymmetries in the empirical analysis of PPP behaviour between South Africa and her main currency exchange partners. Secondly, seeing that the empirical dataset consists of the nominal exchange rates and price level indices; this implies that these time series variables "...incorporate prices of non-traded goods; it is unlikely that their use in an empirical test would produce symmetry and proportionality..." (Macdonald, 1995).

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 results of the Kapetanois and Shin (2006) unit root tests as employed on the time series variables.

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

	test statistics									*threshold estimates*	
	W_{sup}			W_{ave}			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:"***", "**' and '*' denote the 1%, 5% and 10% significance levels respectively. Critical values at 10 percent significance level:

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 the results reported in

Table 1, concerns the threshold estimates which determine the rand value at which the real exchange rate is found to be stationary. Take for instance, the finding of the PPP hypothesis being found to valid only when the computed real effective exchange rate between the rand the US dollar as well as the British pound lies outside the range of \$1=R7.95 and \$1=R9.17 for the dollar and between £1=R7.95 and £1=R9.17 for the pound. Similar inferences can be drawn for the Euro, the Chinese Renminbi and the Japanese Yen with stationary processes being found outside the real exchange rates of €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.

One noteworthy advantage of the performed unit root tests is that they render the time series variables as a regime-switching function of both a unit root process as well as a stationary process. This is important in our empirical analysis, since, on one hand, this can render the stationary portion of the real exchange rates as being in complete compliance with the PPP hypothesis, and on the other hand, it renders the unit root portion of the nominal exchange rate and the differences in the consumer price indices as providing preliminary evidence of PPP cointegration. As is demonstrated in table 1, the results indicate a partial unit root process for the nominal exchange rates and the differences in price indices between South Africa and all her main trading partners. Therefore, the performed unit root tests provide two separate and yet simultaneous evidences of PPP behaviour between South Africa and her main trading partners. 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 or conformity of 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 proceeds to implement 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.

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

	dependent varaible Y	*independent varaible* X	$(c)TAR$				$(c)MTAR$			
			$H_0^{(1)}$	$H_0^{(2)}$	$H_0^{(3)}$	$H_0^{(4)}$	$H_0^{(1)}$	$H_0^{(2)}$	$H_0^{(3)}$	$H_0^{(4)}$
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)	(0.13)*	*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-statistics are reported in ().

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 cannot be rejected for all PPP threshold co-integration residuals. Secondly, we also note that all regressions significantly manage to reject the null hypothesis of no co-integration effects between nominal exchange rates and the differences in domestic and foreign aggregate prices for all threshold regression specifications. This result is of particular importance because it indicates there is a significant PPP relationship between South Africa and her main currency trading partners. Thirdly, we are unable to reject the null hypothesis of symmetric co-integration between the variables for all model specifications. However, we find that for the threshold co-integration relationship between South Africa and each of her currency trading partners, we find at least one significant threshold model (for all trading partners with the sole exception for Euro data). For instance, for the case of South Africa's currency relationship with the US, we are unable to find threshold cointegration using the TAR and MTAR models with zero threshold estimates and, yet we find significant threshold co-integration effects with the c-TAR and c-MTAR models which include a consistent threshold estimate. Similarly, we are able to only establish significant c-MTAR threshold co-integration effects for the UK whereas for both China and Japan we find significant MTAR and c-MTAR threshold effects for the observed data. Lastly, we find that for each of the established threshold co-integration models, we are able to establish significant threshold error correction effects.

Based on the results reported in Table 2 we observe a number interesting phenomenon. In particular, we observe that for all threshold cointegration regressions there exists a smooth (as indicated by a MTAR model) as opposed to an abrupt (as indicated by a TAR model) co-integration and error correction transition mechanism between South African nominal exchange rates and the differences in aggregate prices. An exception is noted for estimates on US data, in which the empirical results reveal the existence of a TAR-TEC regression model in addition to the MTAR-TEC specification. This implies both a smooth as well as an abrupt threshold cointegration between the currency exchange relations of South Africa and the US. Having generally established relevant and significant threshold co-integration and error correction effects for all PPP regression specifications, we proceed to formally estimate the associated TAR-TEC and MTAR-TEC models for the relevant regressions, with the estimation results being reported below in Table 4. For each of the estimated regressions, we provide estimates of threshold value (where applicable), as well as the TAR or MTAR regressions and the associated TEC specification.

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

For US

$c-TAR-TEC$

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

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

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

$c-MTAR-TEC$

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

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

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

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

$MTAR-TEC$

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

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

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

$c-MTAR-TEC$

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

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

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

For JAPAN

$MTAR - TEC$

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

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

$$\Delta\epsilon_t = \underset{(0.60)}{0.00} + \begin{cases} \underset{(0.72)}{-0.26}\Delta\pi_{t-1}^* + \underset{(0.01)^{**}}{0.30}\,\Delta\epsilon_{t-1} - \underset{(0.57)}{0.02}\,\xi_{t-1}I.(\Delta\xi_{t-1} \geq 0) \\ \underset{(0.69)}{-1.26}\Delta\pi_{t-1}^* + \underset{(0.12)^{*}}{0.28}\,\Delta\epsilon_{t-1} - \underset{(0.01)^{***}}{0.08}\,\xi_{t-1}I.(\Delta\xi_{t-1} < 0) \end{cases}$$

$c - MTAR - TEC$

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

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

$$\Delta\epsilon_t = \underset{(0.69)}{0.00} + \begin{cases} \underset{(0.85)}{-0.13}\Delta\pi_{t-1}^* + \underset{(0.01)^{***}}{0.28}\,\Delta\epsilon_{t-1} + \underset{(0.55)}{0.20}\,\xi_{t-1}I.(\Delta\xi_{t-1} \geq 0.02) \\ \underset{(0.64)}{-1.41}\Delta\pi_{t-1}^* + \underset{(0.11)^{*}}{0.28}\,\Delta\epsilon_{t-1} - \underset{(0.00)^{***}}{0.09}\,\xi_{t-1}I.(\Delta\xi_{t-1} < 0.02) \end{cases}$$

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

From the results reported above in Table 3, all estimated regressions indicate significant positive correlations between nominal exchange rates and the difference in aggregate price levels; which is a result consistent with the PPP theoretical hypothesis. In applying Hansen's (2000) conditional least squares (CLS) method to estimate the threshold values for the c-TAR-TEC and c-MTAR-TEC models; we obtain values which lie in the range of between -0.08 and 0.02. These obtained threshold values which govern the regime switching behaviour of the error terms can be considered very reasonable estimates since they all lie relatively close to a value of zero. In particular, for a threshold value close to zero, positive discrepancies from the long-run equilibrium are measured by the absolute value of the coefficient on the error term above the threshold value (i.e. $| \rho_1 |$) whereas negative discrepancies are measured by the absolute value of the coefficient on the error term below the threshold value (i.e. $| \rho_2 |$). Based on the results reported in Table 3, it is evident that the speed of adjustment towards equilibrium is more rapid for positive discrepancies in the case of South African-US and South African-UK relations whereas convergence towards the long-run equilibrium is more rapid for negative discrepancies for the South African-China and the South-Africa-Japan cases.

On the other hand, the estimated error correction coefficients ξ_{t-1}^- and ξ_{t-1}^+respectively provide a measure of 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. In general, 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 Africa-China 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 Japan-South Africa 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 each of estimated threshold cointegration and error correction models as was described in detail in the previous section of this paper. The results of the granger causality tests are reported in Table 4 which show that for all estimated equations; nominal exchange rates (i.e. ϵ_t) are deemed to granger cause aggregate price levels (i.e. π_t^*). This result is in coherence with those obtained in Kholdy and Sohrabian (1990) as well as Schnabl and Baur (2002). Einzig (1935) attributes this finding to the notion that in a system of flexible exchange rates appreciation (depreciation) of a country's currency leads to a decrease (increase) in the general price level because of the impact on domestic activity. Conversely, an appreciation (depreciation) dampens (stimulates) domestic activity, inflation is curbed (accelerated) and the central bank will adapt monetary policy by slowing (accelerating) monetary expansion. Another perspective as presented by Hafer (1989) insinuates that import prices can act as a transmission mechanism from the exchange rate to domestic prices. Furthermore Menon (1995) attributes this behaviour to exchange rates affecting domestic prices through export prices due to incomplete pass through and productivity adjustments. All-in-all, either of these can hold as a suitable explanation in providing a relevant explanation as to the causality results as obtained in our empirical analysis.

Table 4: Granger Causality tests

$country$	$model\ type$	Y	X	$H03$	$H04$	$decision$
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 ().

5 *CONCLUSIONS*

In view of a lack of evidence in 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 gap in the academic paradigm in a two-stage empirical process. In the first stage of our empirical analysis, we examined 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 between South Africa and her main currency trading partners. In further applying the aforementioned unit root tests to nominal exchange rates and price rate differentials; partial evidence of PPP cointegration was preliminary established as the time series were found to partial containing a unit root process.

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 where asymmetrically co-integrated. To this end, the empirical results were able to confirm significant asymmetric cointegration evidence for all South Africa's currency trading partners with the sole exception of the Euro area. In particular, the empirical analysis depicted that negative deviations from the long-run equilibrium are easier to eradicate within the South-African UK and South African-US relations whereas negative deviations from long-run equilibrium are found to persist longer for the South Africa-Chinese and South Africa-Japan cases. However, for all relations exchange rates are the primary mechanism of adjustment toward the long run equilibrium between South Africa and her main trading partners. Having established significant asymmetric cointegration relations between the various PPP relationships, we supplemented this evidence with granger causality tests. Contrary to popular 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, in 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 aggregate price levels can be achieved through stability in exchange rate 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 the purchase foreign exchange from financial markets. In terms of policy implications, our results therefore depict that an exchange rate targeting framework may prove to be a useful avenue for future policy stabilization policies which may be adopted by the SARB. As a by product, the empirical results obtained in this study further supplement 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 viable alternative to the strict pursuing of an inflation-targeting regime which is currently under heavy criticism for being a rather strict monetary policy regime.

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