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Bernard Njindan Iyke[1] and Nicholas M. Odhiambo

Abstract

In this paper, we have estimated the equilibrium real exchange rate for Botswana. We have also reviewed the exchange rate regimes pursued by Botswana from independence to date. The evidence suggests that Botswana operated a fixed exchange rate without adjustable pegs from 1966-1976; a fixed exchange with adjustable pegs from 1976-1980; and a fixed exchange with a currency basket from 1980 to date. From the ARDL bounds testing procedure, we found that the fundamental determinants of the equilibrium real exchange rate in Botswana are: the terms of trade and trade openness. The actual real exchange rate appears to have deviated significantly from the equilibrium exchange rate. Perhaps more worrying is the fact that our estimated speed of adjustment is very slow. This means that significant deviations are not corrected fast enough annually. Policymakers in Botswana are encouraged to pursue policies, which could raise the adjustment parameter, in order to avoid excess misalignments, when going forward.

Keywords: *Real Exchange Rate, Equilibrium Exchange Rate, Botswana*

JEL Classification Code: *C53, F31*

1. Introduction

What approach is appropriate for estimating equilibrium exchange rates? There have been an increasing number of research papers proposing various approaches for estimating the equilibrium exchange rate. Central to this renaissance is the fact that measuring, detecting and

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correcting the real exchange rate disequilibrium is crucial to achieving and maintaining macroeconomic stability (see, for instance, Edwards, 1996). It is well known that real exchange misalignments are not favourable for economic growth (see Rodrik, 2008). Developing countries are even more cautious about this because other factors are already inhibiting their growth prospects. Thus, the real exchange rate remains a crucial variable which has been at the forefront of many policy discussions in developing countries (see Melvin, 1985; Collins, 1996; Papaioannou, 2003, among others).

This paper aims to contribute to the literature on the real exchange rate estimation by attempting to provide the equilibrium real exchange rate for Botswana for the period 1975-2012. We take a fairly clear stance by adopting the fundamental determinants of equilibrium exchange rate approach. This approach basically entails identifying the fundamental macroeconomic factors which influence real exchange rates in theory and in practice. A representative literature for this approach is the optimum currency area (OCA) literature which identifies economic size and trade openness as long-run fundamental determinants of the choice of exchange rate regimes (see Mundell, 1961; and McKinnon, 1963; for detail exposition of the OCA literature). Other studies which also motivate the fundamental determinants approach include Frankel (1995), Tornell and Velasco (1995), Edwards (1996), Berger *et al.* (2000), and Hausmann *et al.* (2001).

Instead of proceeding to estimate the equilibrium exchange rate for Botswana using the Engle-Granger two-step technique or the Johansen technique, as has been the case in the previous studies, we use the autoregressive distributed-lag (ARDL)-bounds testing technique, which is shown to perform well in small samples (see Pesaran *et al.*, 2001). This is another contribution of this paper, in addition to a detailed chronology of exchange rate events that ensued in Botswana

from independence to date. To the best of our knowledge, this study is the first of its kind to expose the exchange rate events and provide a solid estimation of the long-run and short-run movements of the exchange rate in Botswana.

The next section presents the chronology of the exchange rate events in Botswana from independence to date. In section 3, we provide a brief review of the relevant literature on equilibrium real exchange rate estimation. Section 4 discusses the econometric methodology. Section 5 reports the results of our estimations, and section 6 provides the concluding remarks.

2. The Chronology of Exchange Rate Events in Botswana from Independence to Date

Botswana was a member of the Rand Monetary Area (also known as the Common Monetary Area) from Independence in 1966 till 1976. Under the Rand Monetary Area framework, the South African rand was the legal tender in Botswana. The rand was pegged against the British pound and the US dollar (see Masalila and Motshidi, 2003). In 1976, Botswana adopted a fixed exchange regime with adjustable pegs when the Botswana currency, the pula, was introduced as the legal tender. The pula was pegged to the US dollar at 1.15 per dollar in the same year, meaning that the pula was on par with the rand (Bank of Botswana, 2001).

The first significant intervention in the foreign exchange market of Botswana took place in April 1977, when the pula was revalued by 5 per cent, in order to control inflation (Masalila and Motshidi, 2003). Also, in June 1980, the pula was unpegged from the US dollar. A currency basket comprising the rand and the Special Drawing Rights (SDR) was introduced for the pula. This was mainly implemented, in order to minimize the volatility between the pula and the rand. In November of the same year, the pula was revalued by 5 per cent by the Bank of Botswana

which aimed to contain the imported inflation following the depreciation of the pula against the rand. In May 1982, the Bank of Botswana intervened in the foreign exchange market once more. This time, the pula was devalued by 10 per cent, as a measure to resolve the balance of payment crisis of the period 1981-1982 (see Jefferis and Harvey, 1995; Masalila and Motshidi, 2003).

The impact of the rand on the pula seems to be very strong. In February 1984, the rand depreciated significantly against the dollar, due to a foreign debt overhung. The pula depreciated significantly against the dollar as a result. The Bank of Botswana decided to undertake another devaluation exercise, in order to enhance the external competitiveness of the country's export; since the pula had experienced a significant appreciation against the rand. Thus, the pula was devalued by 5 per cent against the rand. In August of the same year, the weight of the rand in the pula basket was adjusted by the authorities, in order to maintain a close relation between the rand and the pula. This was then followed by a devaluation exercise in January 1985, when the pula was devalued by 15 per cent. The goal of this devaluation exercise, according to the Bank of Botswana, was to maintain the external competitiveness of the pula (see Government of Botswana, 1985).

From January 1985 to January 1986, the rand appreciated significantly against the pula. Consequently, the Bank of Botswana reacted by introducing a new pula basket in January 1986, in order to align the pula with the rand. After this intervention, the foreign exchange market was not tampered with by the Bank of Botswana for at least two years. However, in June 1989, the pula was revalued by 5 per cent to contain imported inflation. The new rate did not last long; as the authorities decided to devalue the rate by 5 per cent in August 1990, and again in August 1991, in a bid to enhance the external competitiveness of the pula (see Jefferis and Harvey,

1995). In June 1994, the monetary authority of Botswana decided to adjust the weight composition of the pula basket (which contains the SDR and the rand), but only on technical grounds. According to Masalila and Motshidisi (2003), the rand has been assigned greater weight due to the significant proportion of economic activities, which are rand-denominated. From 1994 to date, Botswana operated a fixed exchange regime, with the pula pegged to a basket of currencies.

3. A Review of Related Literature

A large body of empirical studies has provided alternative measures for the estimation of the equilibrium real exchange rate. The curious reader is referred to MacDonald (1995), Rogoff (1996), and Hinkle and Montiel (1999) for surveys on the relatively old studies. For the most recent studies, the reader could refer to Lee *et al.* (2008). A large number of these studies have been very critical on the choice of the purchasing power parity (PPP) procedure for estimating the equilibrium real exchange rate. MacDonald and Ricci (2003) argue, for example, that the PPP model suffers greatly from its slow mean-reverting property to a constant level, which is its implied long-run equilibrium assumption. Quite a number of these studies argue that a better way to estimate the equilibrium exchange rate it to identify fundamental macroeconomic drivers of the real exchange rate. The recommendation is that an appropriate model be fitted, within which the dynamic response of the real exchange rate to shocks from these macroeconomic variables can be examined (see Montiel, 1999).

In the classical setting, the equilibrium exchange rate for a country could be estimated by building and simulating an empirical dynamic macroeconomic model using parametric

calibrations and data suitable for that country. In the literature, this approach is known as the general equilibrium approach. Montiel (1997) argues that Monte Carlo simulations on the specific policy paths and the exogenous determinants result in consistent steady-state values for the real exchange rate, if an appropriate macroeconomic model is built for a country. Studies that have employed this procedure to examine the exchange rate equilibrium include Clark *et al.* (1994), Williamson (1994), and Stein *et al.* (1995).

Another approach to estimate the equilibrium exchange rate is the single equation models and partial equilibrium models approach. The common technique under this approach is the 'trade elasticities' technique, which is a partial equilibrium technique (see Montiel, 1999). Driver and Wren-Lewis (1997), for instance, have used this technique to estimate the real exchange rate equilibrium for the G7 countries. Ghei and Pritchett (1999) also used this technique in their study. Some researchers have, instead, employed reduced form equations to estimate the equilibrium real exchange rate. This technique has often required that the fluctuation components of the real exchange rate be removed. A classic study, which employed this technique, was that of Edwards (1994) for a panel of 12 developing countries.

The most recent approach draws econometric techniques which sought to gain robust parameter estimates from the non-stationary properties of the macroeconomic determinants of the real exchange rate. This is known as the 'cointegration approach' for estimating the equilibrium real exchange rate. Various studies have used the cointegration approach in the literature. The key studies are by Elbadawi and Soto (1994), Cardenas (1997), Montiel (1997), Loayza and Lopez (1997), Aron *et al.* (2000), MacDonald and Ricci (2003), and Iman and Minoiu (2011).

This paper favours the cointegration approach. The main motivation is that the cointegration approach is able to capture any short-run fluctuations in the movement of the real exchange rate, and thus, represents what is, indeed, observed in reality. The studies which employ the fundamental approach have identified a number of macroeconomic drivers of the real exchange rate. The most common macroeconomic variables are: world commodity prices (see Cashin *et al.*, 2002; Chen and Rogoff, 2002; MacDonald, 2002; MacDonald and Ricci, 2003), terms of trade (see Montiel, 1997; Goldfajn and Valdes, 1999; Imam and Minoiu, 2011), government consumption or spending (see De Gregorio *et al.*, 1994; Montiel, 1997; MacDonald and Ricci, 2003; Imam and Minoiu, 2011), size of the net foreign assets (see Lane and Milesti-Ferretti, 2000; MacDonald and Ricci, 2003; Imam and Minoiu, 2011), and trade openness (see Montiel, 1997; MacDonald and Ricci, 2003; Imam and Minoiu, 2011). These are the variables employed in this study. One crucial variable, the growth of the world economy, is included in the estimation. This study is, perhaps, the first to include this variable. The inclusion of the growth of the world economy in this estimation is motivated by the fact that real exchange rates, especially in more open economies, are known to respond to the performance of the global economy.

4. The Econometric Technique

This paper uses the autoregressive distributed lag (ARDL) bounds-testing procedure advanced by Pesaran *et al.* (1996), Pesaran and Shin (1999), and Pesaran *et al.* (2001) to investigate the long-run relationships between the real exchange rate and its fundamental determinants. This procedure is preferred to other procedures because: (i) It can be applied, whether the time series

are integrated of order zero, one, or a mixture of both; and (ii) it has better finite sample properties.[2]

The ARDL bounds-testing procedure requires that we employ the following general $ARDL$ equation, in terms of our variables:

$$\Delta lnREER_t = \gamma_0 + \gamma_1 t + \sum_{i=1}^{k} \gamma_{1i}\, \Delta lnREER_{t-i} + \sum_{i=0}^{k} \gamma_{2i}\, \Delta lnTOT_{t-i} + \sum_{i=0}^{k} \gamma_{3i}\, \Delta lnOPEN_{t-i} + \sum_{i=0}^{k} \gamma_{4i}\, \Delta lnGCR_{t-i} + \sum_{i=0}^{k} \gamma_{5i}\, \Delta lnNFAR_{t-i} + \sum_{i=0}^{k} \gamma_{6i}\, \Delta lnRCPI_{t-i} + \sum_{i=0}^{k} \gamma_{7i}\, \Delta lnWGDPG_{t-i} + \delta_1 lnREER_{t-1} + \delta_2 lnTOT_{t-1} + \delta_3 lnOPEN_{t-1} + \delta_4 lnGCR_{t-1} + \delta_5 lnNFAR_{t-1} + \delta_6 lnRCPI_{t-1} + \delta_7 lnWGDG_{t-1} + \mu_t \quad (1)$$

Where μ, γ, and δ are, respectively, the white-noise error term, the short-run coefficients, and the long-run coefficients of the model; and Δ is the first difference operator. In addition, *t* denotes time period; k is the maximum number of lags in the model.

From *Equation (1)*, there exist cointegrating relationships between the series, if at least one of the δs is significantly different from zero. Otherwise, we could reject the evidence of cointegration between real exchange rate, in particular, and the fundamental determinants; since any other possible evidence of cointegration would not be relevant to our study.

We note that the ARDL bounds-testing procedure for cointegrating relationships follows a non-standard asymptotic *F*-distribution. Two sets of critical values have been constructed by Pesaran *et al.* (2001) under this null hypothesis. The first set of critical values is constructed under the assumption that variables in the ARDL model are integrated of order zero, *I(0)*. The second set

[2] See Pesaran *et al.* (1996), Pesaran and Shin (1999), and Pesaran *et al.* (2001), for extensive discussion.

of critical values is constructed under the assumption that variables in the model are integrated of order one, *I(1)*. We do not reject the null hypothesis of no cointegration relationships when the *F*-statistic falls below the lower-bound values. Similarly, we reject the null hypothesis of no cointegration when the calculated *F*-statistic is greater than the upper-bound values. However, the test is inconclusive, when the *F*-statistic falls between the lower and upper bounds.

5. The Estimated Equilibrium Real Exchange Rate for Botswana

The fundamental determinants of the equilibrium real exchange rate in Botswana were extracted, by drawing insight from the empirical literature we have discussed earlier. We extracted the real effective exchange rate (REER), the ratio of government consumption to GDP (GCR), trade openness (OPEN)[3], terms of trade (TOT)[4], net foreign assets to GDP (NFAR), and the world economic growth (WGDPG) from the World Development Indicators (WDI), 2014. The data on real commodity price index (RCPI)[5] were extracted from the Global Economic Monitor (GEM) Commodities, 2014.

As a preliminary analysis, prior to estimating the equilibrium real exchange rate for Botswana, we examined the stationary properties of the variables. We employed the Dickey-Fuller Generalized Least Squares (DF-GLS) and the Ng-Perron tests, proposed by Elliot *et al.* (1996), and Ng and Perron (2001), respectively. Our choice was influenced by the fact that the canonical tests for stationarity (such as the ADF and PP tests) were found to reject the null hypothesis of

[3] This ratio is measured as the sum of exports and imports of goods and services to GDP (in 2005 constant US$).

[4] This ratio is measured as imports divided by exports of goods and services (see Obstfeld and Rogoff, 1996).

[5] We calculated this index by principal component analysis for prices of sugar, beef, and copper which are the major exports of Botswana in 2013, according the Central Statistics Office of Botswana.

unit root, when the time series under consideration has a large and negative moving average (MA) component, even when there is a unit root (see Schwert, 1986; Caner and Killian, 2001). Elliot *et al.* (1996), and Ng and Perron (2001) demonstrated, respectively, that the DF-GLS and the Ng-Perron tests have substantially higher power, even when the root of the time series is closer to unity.

All the variables employed in the equilibrium real exchange rate model for Botswana were found to be stationary at the 1 per cent level of significance, when differenced once, except WGDPG, which was stationary at level. Trend was not included in each of the tests, as the graph of the variables[6] did not indicate any significant trending phenomenon. Besides, lag choice was limited to a maximum of 4, using the Modified Akaike Information Criterion (MAIC). This choice was informed by the small size of the dataset.

Table 1: Tests for Unit Roots of the Variables

Variable	DF-GLS t-statistic	Ng-Perron MZ_a-statistic
LNREER	-1.3981	-3.3923
ΔLNREER	-5.3513***	-17.7961***
LNGCR	-1.3237	-2.8071
ΔLNGCR	-3.8553***	-14.7275***
LNNFAR	-1.4194	-0.1949
ΔLNNFAR	-4.6018***	-15.4296***
LNOPEN	-1.2494	-3.1824
ΔLNOPEN	-5.3724***	-17.5645***
LNRCPI	-0.8014	-2.2590
ΔLNRCPI	-5.2477***	-17.7121***
LNTOT	-1.5796	-4.0572
ΔLNTOT	-4.8719***	-17.2126***
WGDPG	-3.9680***	-15.1916***

Note:

(i) The DF-GLS statistic is based on MacKinnon (1996)

(ii) The Ng-Perron statistic is based on Ng and Perron (2001, Table 1)

[6] See Appendix A, Figure 1 for the graphs.

(iii) *** denotes significance at 1%.

(iv) Δ is a first difference operator.

As the next step, we established the optimal lag of the ARDL model using the AIC. The optimal lag chosen was at most 2. We proceeded to perform the ARDL bounds test on the real exchange rate equation and found that there was evidence of cointegrating relationships between the real exchange rate and its fundamental determinants. The calculated *F*-statistic (5.61) was found to be greater than the upper bound critical value (4.43) at the 1 per cent level of significance. The result for the cointegration test is presented in Table 2.

Table 2: ARDL Bounds Test for Cointegration

Dependent Variable	Function	F-statistic				
lnREER	*lnREER(lnGCR, lnOPEN, lnTOT, lnNFAR, RCPI, WGDPG)*	5.61***				
Asymptotic critical values for unrestricted intercept and no trend reported from Table CI(iii) p. 300 of Pesaran et al. (2001)	1%	1%	5%	5%	10%	10%
	I(0)	I(1)	I(0)	I(1)	I(0)	I(1)
	3.15	4.43	2.45	3.61	2.12	3.23

Note: *** denotes significance at the 1% level.

Having established that the series are cointegrated, we proceeded to estimate the equilibrium real exchange rate in two stages. In the first stage, we estimated the general short- and long-run models (which are not shown here). In the second stage, we sequentially deleted the variables which were not statistically significant in the long run until we arrived at parsimonious short- and long-run models. The results of the parsimonious models are presented below:

$$lnREER_t = 4.674 - .298lnOPEN_t - .425lnTOT_t \quad (2)$$

$$[162.976^{***}] \quad [-1.708^{*}] \quad [-3.305^{***}]$$

$$\Delta lnREER_t = -.295ECM_{t-1} - .088\Delta lnOPEN_t - .069\Delta lnTOT_t + .131\Delta lnTOT_{t-1} \quad (3)$$

$$[-3.342^{***}] \quad [-1.929^{*}] \quad [-1.637] \quad [2.893^{***}]$$

Equations (2) and *(3)* represent, respectively, the long- and short-run relationships between the real exchange rate and its fundamental determinants – in this case, trade openness and terms of trade. Our estimates show that terms of trade and trade openness are the only significant fundamental determinants of the equilibrium real exchange rate in Botswana.[7] The results show that a 1 per cent increment in trade openness results in approximately 0.30 per cent depreciation in the real exchange rate in the long run, given that the terms of trade remains constant.[8] In addition, we found that an increment in terms of trade by 1 per cent leads to a depreciation of the real exchange rate by 0.43 per cent in the long-run, given that trade openness remains the constant.[9] Finally, the results show that approximately 29.5 per cent of the disequilibrium is corrected each year. The speed of adjustment, when the real exchange rate, trade openness, and the terms of trade variables drift apart was considerably slow.

[7] The t-statistics are shown in the parenthesis; *, and *** denote 10% and 1% significance level, respectively.

[8] Lee et al. (2008), and Imam and Minoiu (2011) found similar results in their studies. However, Montiel (1997), and MacDonald and Ricci (2003) found trade openness to have a positive influence on the equilibrium real effective exchange rate in the long run.

[9] Montiel (1997), Lee et al. (2008), and Imam and Minoiu (2011) all found terms of trade to exert a positive influence on real exchange rate.

The estimated equilibrium passed the recommended tests for structural stability, namely: (i) The cumulative sum of the recursive residuals (CUSUM); and (ii) the cumulative sum of the squares of recursive residuals (CUSUMSQ), proposed by Brown *et al.* (1975). Figure 1 and Figure 2 show the plotted graphs of CUSUM and CUSUMSQ.

Figure 1: Plot of Cumulative Sum of Recursive Residuals (CUSUM)

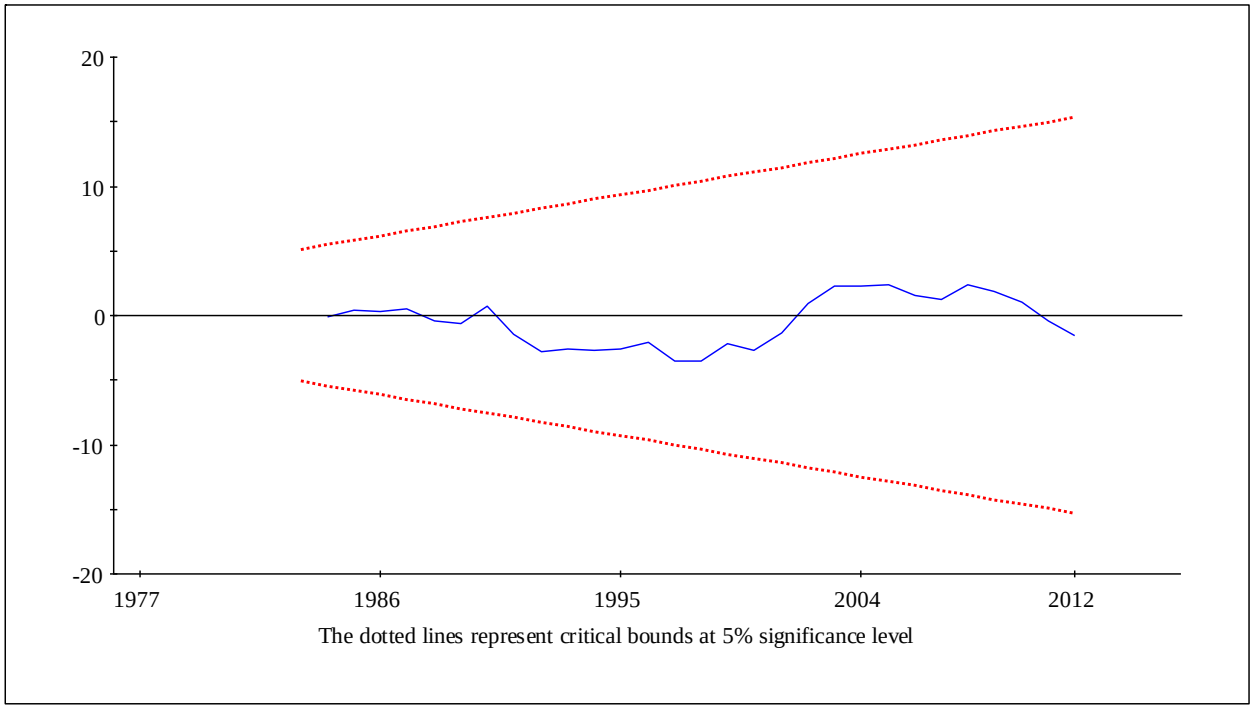
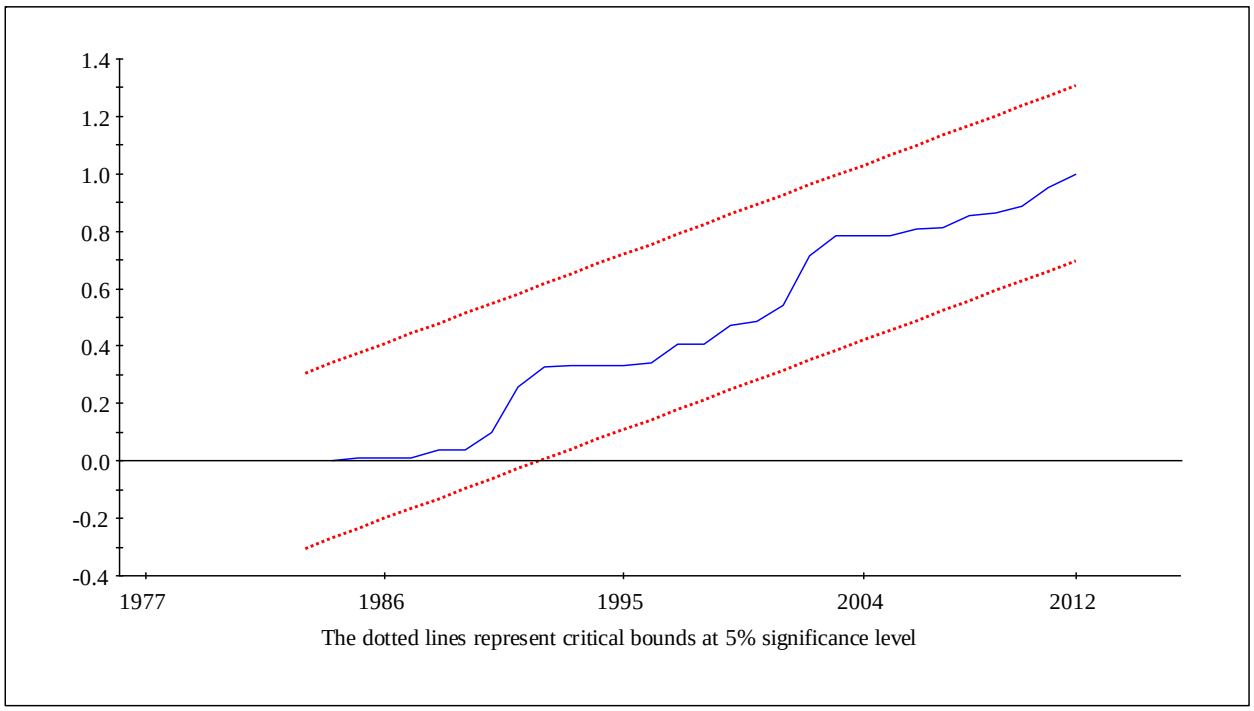


Figure 2: Plot of Cumulative Sum of Squares of Recursive Residuals (CUSUMSQ)



We therefore calculated the equilibrium real exchange rate for Botswana defined as the measure of real exchange rate that is consistent with the long-run values of the fundamental determinants. To neutralize the effects of noise or the cyclical components of trade openness and the terms of trade on the estimated equilibrium real exchange rate, we employed the Hodrick-Prescott (HP) filter[10] time series decomposition technique.

Figures 3 and 4 show the plotted equilibrium real exchange rate (unfiltered and filtered, respectively) against the actual real exchange rate. The equilibrium gap, defined as the difference between the actual real exchange rate and the filtered equilibrium rate, is plotted in Figure 5. The real exchange rate has drifted from the estimated equilibrium exchange rate for most of the sample period, except between 1986 and 1987, and between 1993 and 1996, as Figure 3 shows.

[10] See Appendix B for notes on the Hodrick-Prescott filter which was advanced by Hodrick and Prescott (1997).

The actual rate was highly misaligned upwards from 1980 to 1986. The real exchange rate also experienced rapid depreciation between 1983 and 1988, and a rapid appreciation between 1998 and 2003 (see Figure 3).

The developments in terms of trade and trade openness were not entirely clear-cut. Thus, it becomes difficult to account for the earlier depreciation (1983-1988), and the latter appreciation (1998-2003). Throughout the sample period, the largest equilibrium gap was recorded in 1984 – a gap of 15 per cent. Again, this gap could not be properly explained by the size of the trade openness or by the terms of trade. In fact, the sizes of these two ratios in 1984 were not the highest recorded during the sample period. We note that 29.5 per cent of the misalignment of the actual real exchange rate from its equilibrium level (i.e. the gap) has been corrected, as pointed out earlier.

Figure 3: Plot of Actual against Equilibrium Exchange Rate

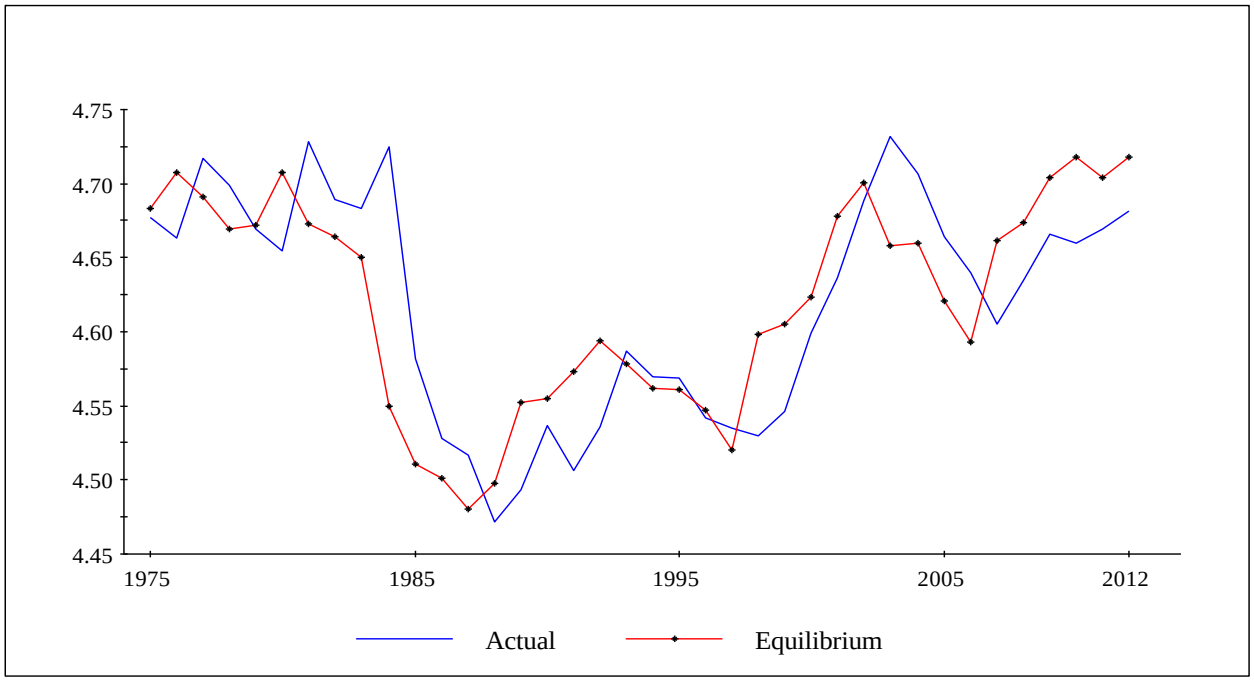


Figure 4: Plot of Actual against the Hodrick-Prescott filtered Equilibrium Exchange Rate

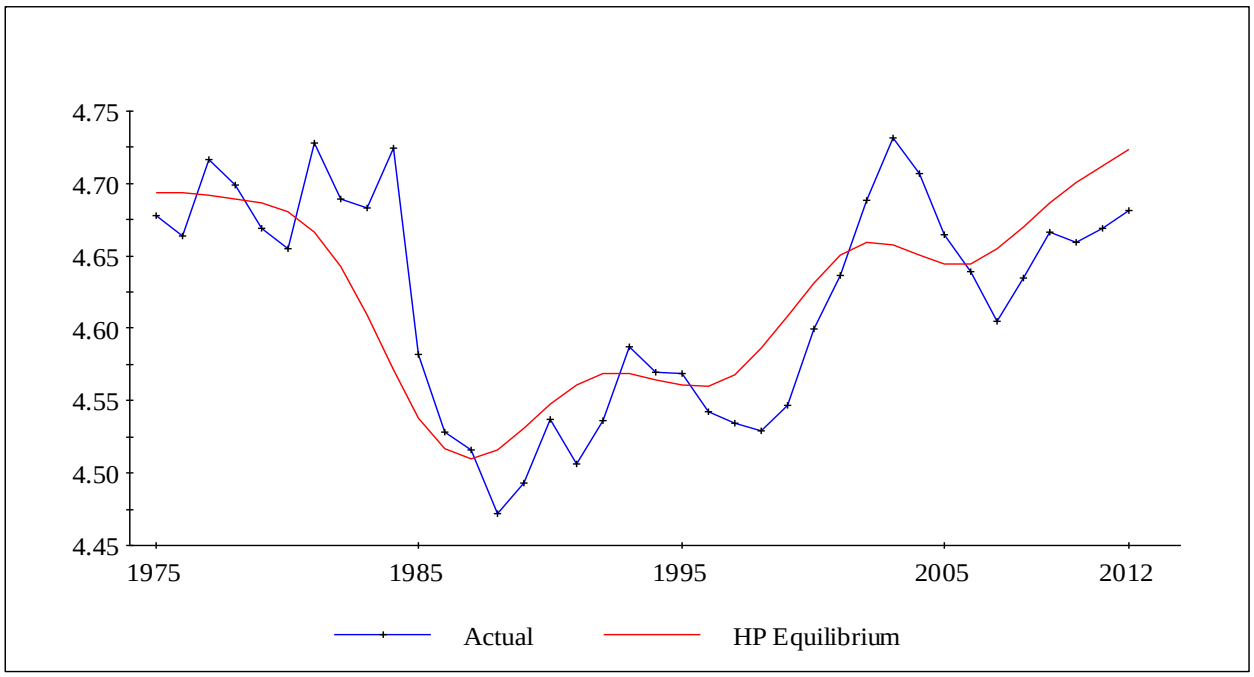
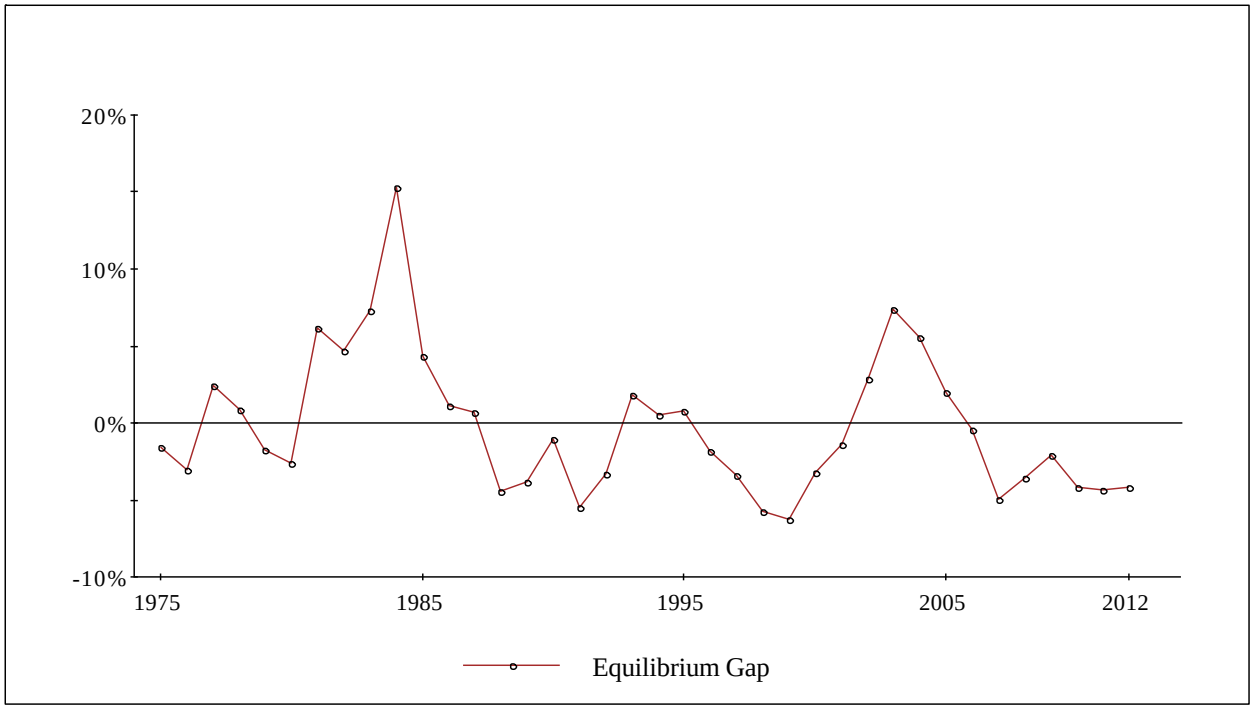


Figure 5: Plot of Equilibrium Gap (Actual minus HP filtered Equilibrium)



6. Concluding Remarks

The exchange rate has remained one of the most widely discussed variables in the global economy. Central to the dominance of this macroeconomic variable is the fact that frequent exchange rate misalignments are generally seen as detrimental to current and capital account balance, and to the accumulation of external reserves. Moreover, the real exchange rate has strong links with other crucial macroeconomic variables, such as inflation and the real interest rate. Thus, policymakers cannot undertake any meaningful short- or long-term projections without incorporating this variable. One of the widely recognized ways of tracking the real exchange rate is by examining its steady-state or equilibrium properties. In this way, policy-makers could determine whether the actual real exchange rate has been moving along the desired path. In the past, researchers have employed the PPP procedure to track the movement of the real exchange rate over time.

However, most recent studies have found the PPP procedure to be inappropriate, due to its slow mean-reverting property. The widely favoured method in the recent literature is the fundamental equilibrium approach. In this paper, we estimated the equilibrium real exchange rate for these countries by following the fundamental equilibrium real exchange rate approach. Unlike most of the previous studies which employed the Johansen, Engle two-step, and Phillips-Ouliaris cointegration procedures, which are known to suffer greatly in small samples, we employed the ARDL bounds-testing procedure, which has better small-sample properties. We also explored the chronology of exchange rate events in Botswana since independence. The evidence suggested that Botswana operated a fixed exchange without adjustable pegs from 1966-1976; a fixed exchange rate with adjustable pegs from 1976-1980; and a fixed exchange rate with a currency basket from 1980 to date. From the ARDL bounds-testing procedure, we found that the

fundamental determinants of the equilibrium real exchange rate in Botswana were: terms of trade and trade openness. The actual real exchange rate appears to have deviated significantly from the equilibrium exchange rate in Botswana. Finally, the speed of adjustment, when the actual real exchange rate deviates from its equilibrium level, is also considerably slower in Botswana.

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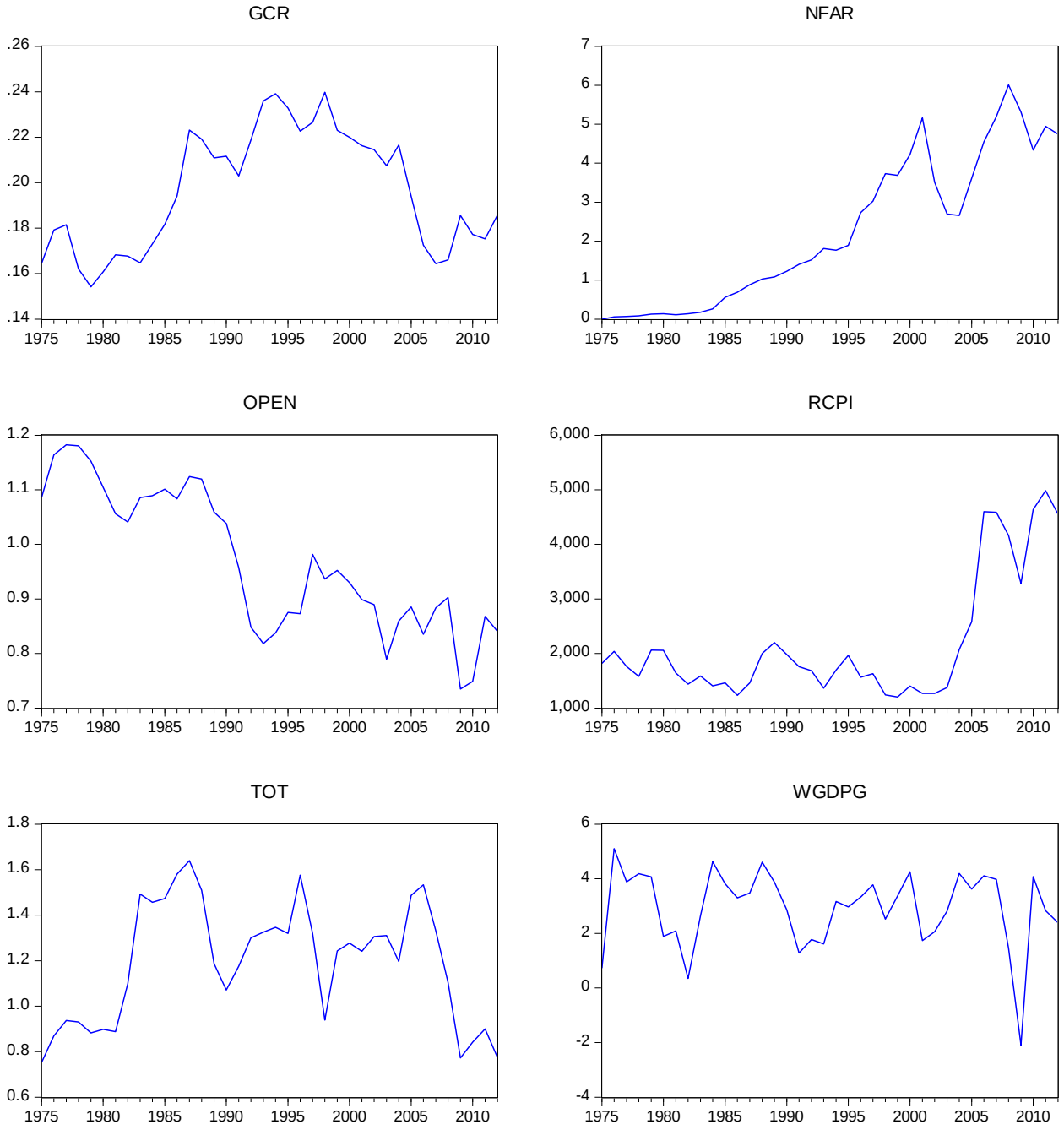
APPENDIX A

Figure 1: The Fundamental Determinants of Real Effective Exchange Rate in Botswana (1975-2012)



Source: Constructed by Authors from WDI (2014) and GEM Commodities (2014)

APPENDIX B

Some Notes on the Hodrick-Prescott Filter

The Hodrick-Prescott (HP) filter or decomposition is a technique employed in macroeconomics and macroeconometrics to decompose a time series into cyclical and trend components. The technique was proposed by Whittaker (1923). But it was not until the seminal paper of Hodrick and Prescott (1997), when the HP-filter gained immense recognition. The main importance of the HP-filter lies in its ability to provide a smooth-curve representation of a time series which is susceptible to long-run impacts than cyclical fluctuations.

The HP-filter is derived on the basis that a time series, say x_t , could be decomposed into a trend (τ_t) and cyclical component (c_t). Assuming, from this intuition, that $x_t = \tau_t + c_t + \mu_t$, where μ_t is the error term of the time series at period *t.* Then, there exist a positive value of a multiplier λ, such that τ solves the minimization problem:

$$\min_{\tau}\left(\sum_{t=1}^{T}(x_t-\tau_t)^2+\lambda\sum_{t=2}^{T-1}[(\tau_{t+1}-\tau_t)-(\tau_t-\tau_{t-1})]^2\right)$$

According to Hodrick and Prescott (1997), the sum of the squared deviations $d_t = x_t - \tau_t$ penalizes the short-run fluctuations in the time series. The second term is a multiple of the multiplier (λ) and the sum of squares of the second differences in the trend component of the series. This term penalizes deviations in the growth of the trend component of the time series. Higher values of λ entails higher penalties. For quarterly data, Hodrick and Prescott (1997) suggest that λ=1600 should be chosen. Ravn and Uhlig (2002) proposed we choose $\lambda = 6.25$ and 129600 for annual and monthly data, respectively.