

University of KwaZulu Natal

THE IMPACT OF EXCHANGE RATE MISALIGNMENTS ON ECONOMIC GROWTH
OF THE SOUTH AFRICAN CUSTOMS UNION.

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Abstract

The dissertation examines the impact of exchange rate misalignments on the economic growth of five countries in the SACU region: South Africa, Lesotho, Namibia, Swaziland and Botswana, using annual data from 1995 to 2012. First and second generation unit root tests are used in order to take into account the existence of structural breaks and cross-sectional dependence. After determining the existence of a cointegration relationship using both Pedroni (2004) and Westerlund (2007) tests, exchange rate misalignments are computed as a deviation of exchange rates from their long-run determinants; estimated using the Pesaran et al.(1997; 1999) mean-group and pooled mean-group. We found that by using the mean-group estimator, the different currencies are overvalued as suggested by Asfaha and Huda (2002) and Saayman (2007), for both the South African and Botswanan currency. Focusing on the results from the mean group, as this estimator is efficient in the presence of cross-sectional dependence in the data, we assessed the impact of misalignment on economic growth using the system-GMM due to the existence of autocorrelation and endogeneity. We found that exchange rate misalignments are not significant in explaining economic growth, even when controlling for terms of trade and openness.

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Chapter 1

1.0 Introduction

This dissertation looks at the impact of exchange rate misalignments on economic growth for South Africa, Lesotho, Namibia, Swaziland and Botswana, given the theory that exchange rate could stimulate economic growth. The dissertation also investigates the existence of policy coordination among these countries which is a condition of monetary union.

Exchange rate misalignment is a key variable when predicting future exchange rate shifts and the need of adjustment of exchange rates among countries with less flexible exchange rate regimes. Sustained exchange rate overvaluation may constitute a warning sign of adjustment of relative prices and a possible decline in the aggregate growth rate of the economy, and by the same token, the exchange rate misalignment can be used to influence the performance of the economy (Magyari, 2008).

Researchers, such as Collins et al, (1997), Aguirre et al., (2006), and Rodrick, (2007) have indicated the existence of a negative relationship between exchange rate misalignment and economic growth, implying that currency undervaluation will spur on growth. There also exists some conflicting results, such as those of Magyari (2008).

A common currency requires the existence of one central bank. If each country has its own central bank controlling money supply, there needs to be cooperation between the banks. Under a common central bank, expansionary or contractionary monetary policy will have an impact on all member countries.

The optimum currency area is characterised by a fixed world price mechanism that is based on a floating real exchange rate, which remains fairly stable even if there is a speculative demand Mongelli (2002). The dynamic equilibrium nature of exchange rates allows reversible movements on exports and imports in competing industries, in an optimum currency area emerging risks of exchange rate is not considerable expensive, cost can be covered at a low cost possible (Mundell, 1961).

The central bank does not engage in monopolistic speculation actions, or optimum currency area permit protection for debtors and creditors which stabilises the economy thereby maintaining long-term flow of capital movement. In recent years, in many geographic locations where countries are within close proximity, the optimum currency area has led to

regional integration with the main goal being to unify the region and create a free trade zone (Kamar and Naceur, 2007). Regional integration leads to an increase in interdependence between countries which may in turn lead to contagious crisis outbreaks. To avoid this problem, member states have to coordinate and harmonize their exchange rates and economic policies, as the real exchange rate is an important measure for assessing a country's competitiveness (see Rodrick 2008; Kamar and Naceur, 2007).

This study examines the impact of exchange rate misalignment on economic growth of Southern African Custom Unions (SACU) countries from 1995 to 2012. The first step compute the real equilibrium exchange rate (REER) using the Behavioural Equilibrium Exchange Rate (BEER) approach advocated by MacDonald (1997) and Clark and MacDonald(1998). As in Magyari (2008), we use the degree of openness, terms of trade and government consumption, a proxy for the Balassa-Samuelson effect; as the determinants of real exchange rate. The BEER is estimated using panel data cointegration technique. The existence of a cointegration relationship implies the existence of some policy coordination among the countries. Misalignments are then measured as deviations of observed exchange rates from the estimated REER; the latter being constructed using the long-run estimates from the cointegrating equation, and the detrended determinants (detrended using the HP filter).

The impact of misalignments on economic growth is estimated using system GMM, due to the existence of endogeneity. However, we have used the misalignment computed after the mean group, as this estimator is efficient compared to the pool-mean group when cross-sectional dependence exists across units.

There exists a cointegration relationship among the variables implying the existence of some form of macroeconomic coordination between the countries under study. The exchange rate misalignments computed using the mean group estimator indicates that all the currencies are overvalued. However, we have not found the existence of a significant correlation between exchange rate misalignment and economic growth.

According to MacDonald (2000) when exchange rate equilibrium is constructed, determinants of real exchange rate may be undermined. It might also ignore relative activity levels and net foreign assets position. Besides, there are various other methodologies that can be used when one constructs the equilibrium exchange rate. Some of these methodologies are the Fundamental Equilibrium Exchange rate of Williamson (1994) and the Natural Real Exchange Rate (NATREX). The estimated REER may therefore depend on the methodology

used. This study does not take also into account the possibility of non-linearity that may exist first between the real exchange rates and its determinants and second, between the variables entering the growth regression. These are some of the limitations of this study.

1.1 Research Question and Objectives

The aims of this study are:

- To establish the existence of macroeconomic policy coordination between the countries;
- To compute the exchange rate misalignments in order to determine if currencies are overvalued or undervalued;
- To analyse the impact of exchange rate misalignments on economic growth.

Hypothesis of the study

Research Questions

- Is there macroeconomic policy coordination among the countries under study?
- Does currency undervaluation or evaluation stimulate economic growth?

Chapter 2

2.0. Literature Review

2.1 Policy Coordination For Regional Integration

Internationally, regional economic integration has become a common phenomenon in recent years. The order of integration begins with weak integration to strong integration, to a free trade zone, a custom union, common market, an economic union and lastly a complete political union.

International trade theory argues that free trade between economies will allow countries to focus on producing goods and services efficiently. Additionally, free trade between countries results in economic growth and brings about foreign direct investment (FDI) inflow through the transfer of technology and skilled labour (Asfaha and Huda, 2002). Regional integration leads to rising interdependence between member states, and is more likely to cause a crisis especially if there is a lack of policy coordination. This was witnessed in the European Union in 1992, and in Latin America in 1994 (Baig and Goldfajn, 1998), and in the last global financial meltdown. To prevent such situations, or to reduce the possibility of running into an exchange rate crisis, member states harmonize their exchange rates and monetary policies (Krugman, 2001). The most advanced stage of coordination between countries is achieved when member states enter into a currency union.

Terra and Valladares (2003) define exchange rate misalignment as deviations occurring around a long-run, real exchange rate equilibrium relationship, and likely to be associated with regime switching thereby allowing for these misalignments to be interpreted as depreciation or appreciation phenomenon. Devarajan (1997) estimates real exchange rate misalignment through using a nominal exchange rate that takes into account price level differences across countries. Aguirre and Calderon, (2005) used bilateral nominal exchange rate for the current period multiplied by domestic price level in the current period divided by foreign current price level.

An increase in regional economic dependence normally generates a need for a regional multicountry currency area, except in situations where it is believed that one country always describes the optimum currency area, implying that political limits and currency borders do not need to be reconciled (Alesina, 2003). There are a series of stages to be followed when creating a single currency among member states, and there should be a period when different

currencies are freely exchanged at a constant rate. This is succeeded by a monetary union with a single currency and a single exchange rate, a single monetary market and unrestricted currency movement and deposit at a constant rate (Kamar and Naceur, 2007). When member states are in a monetary union and banking policies are the same, and there is no central bank managing a pool of foreign exchange reserves, financial market integration leads to the convergence of member states' economies (Buiters, 2000).

Cooper and Kempf's (2003) model used to implement a single currency provides a very favourable view. With all sovereign countries entering into a single currency requiring strong international commitment, there is no single monetary authority. As has been seen with Spanish Mercado Común del Cono Sur (MERCOSUR), a Latin American regional union, the lack of a real exchange rate and coordination in a regional monetary union, which is accompanied by trade barriers leads to the economic turmoil evident in Brazil and Argentina (Kamar and Naceur, 2007). Monetary policy implementation must be coordinated in all countries seeking to use a common currency because their policies should have a similar impact on the exchange rate, whereas if member states have different monetary frameworks, the result is different impacts on real exchange rate (Frankel 1999).

It is important to measure the extent to which monetary policy affects budget deficits, government expenditure, and trade policy on exchange rate movements on all member states to be able to evaluate if these effects of monetary policy have similar outcomes in all member states (Bayoumi and Mauro, 2001). If it is determined that the effects of monetary policy are similar in all member countries, the successful launch of a new currency among member countries could be expected. Alternatively, if it is determined that the effects of monetary policy on all member states are different, especially policies affecting exchange rate behaviour in each country, it is likely that coordination is not sufficient and hence the launch of a common currency is likely to run into problems. This is a situation where there is still a need for a macroeconomic policy harmonization (Kamar and Naceur, 2007).

The real exchange rate defines a relationship between national and international currency prices. When there are changes in high powered money which causes price levels to change and to be different from the price levels prevailing in the rest of the world, the real exchange rate also changes from that of the rest of the world (Sungur, 2004). Edward (1988) argues that the real exchange rate measures the relative prices of goods and services, precisely. Stabilisation is not compatible with fixed exchange rates in relation to free capital movements

(Christiano et al., 1999). The real exchange rates of countries in the GCC which is fixed against the US dollar, have a monetary policy that is connected to the US monetary policy to minimize the effect of interest rates on the real exchange rate behaviour (Laabas and Limam, 2002).

In the region of GCC countries, which usually experience capital flows as a result of increases in the price of oil, it is reasonable to expect an increase in the stock of net foreign assets, thereby leading to an increase in money supply. This leads to a fall in interest rates resulting in a rise in the demand for money, and giving further rise to the supply of money due to an increase in oil price in the Gulf Cooperation Council countries. This increase in the supply of money creates inflationary pressure because money growth is in double digits for Qatar and UAE, which once went as high as 34% (Laabas and Limam, 2002). Kamar and Naceur, (2007) found on many occasions that money growth in the Gulf Cooperation Council countries converged, and on many occasions in the same period of analysis countries have experienced disparities. Kamar and Naceur, (2007) found that Qatar is not correlated with other countries, and Oman was negatively correlated with only one member country, the UAE. On the other hand EAU was also correlated with Saudi Arabia, and Saudi Arabia was correlated with Kuwait and had a strong positive correlation with Bahrain. They also determined that Kuwait started at a higher level but converged rapidly over a period of two years. Kuwait experienced high interest rates, which later declined to the level of its regional partners (Ramanathan, 2007; Setser and Ziemba, 2009).

The method used to finance a budget deficit is important for determining if inflationary pressure is likely to emerge. If it is financed through taxation or by means of internal borrowing, an economy is exposed to the risk of experiencing depressed private spending offsetting a rise in government spending. If a budget deficit is financed through borrowing from the rest of the world, inflationary pressure may be reduced by increasing the supply of imports. Any form of monetization of debt to GDP increases the degree of inflationary pressure (Sala-i-Martin and Sachs, 1991). But in the Gulf Cooperation Council, a country's fiscal deficit does not explain their economic situation, the best explanation for this economic event is budget surplus, which can also explain their inflationary pressure. In addition, all funds needed in the Gulf Cooperation Council are financed mainly through oil reserves instead of borrowing or assets' sales (Setser and Ziemba, 2009).

Budget surplus is a result of oil revenue, not because of increased taxes or reduction of government spending; a decrease in net domestic assets is counteracted by an increase in net foreign assets (Sala-i-Martin and Sachs, 1991). In the Gulf Cooperation Council net foreign assets are more important, and can result in even more high powered money causing inflationary pressure. These countries rapidly convert fiscal surplus into government expenditure thereby increasing aggregate demand and offsetting demand-pull inflation.

However, increases in government expenditure can lead to an increase in public servants' wages, making the private sector, which is competing for employees, increase wage offers and consequently push up inflation. The literature indicates that a negative relationship existing in the budget's balance in the long-run, means an increase in the budget and results in real exchange rate appreciation Kamar and Naceur, (2007).

Common currency allows liquidity service provided by the central bank to circulate over a large geographic reach as preferred method of payment, store of value and unit of account, common currency give way to clear price transparency thereby preventing price discrimination and encouraging competition. Additionally price stability is realised and there is access to a wider and more transparent financial market, which promotes external financing, which mostly benefits countries that have historically been experiencing inflationary pressure. This is because in the economic and monetary union there are plausible anti-inflationary policies Mongelli (2002). Capital mobility rises because of the increasing degree of openness within countries, resulting in a higher likelihood of fluctuations in capital flow. Economic theory is unclear about the exact effect of commercial liberalization. For Gulf Cooperation Council countries, when there is an increase in the rest of the world's price of exports, it means for them an increase in the price of oil, which leads to an increase in capital inflow (Setser and Ziemba, 2009).

Magyari (2008) argues that the real exchange rate is an important macroeconomic policy element, mostly for developing states. Here it is used to estimate fluctuations of the future exchange rate for those countries with flexible exchange rates and to determine the need to move exchange rates among countries with less floating regimes. If the exchange rate is sustained, overvaluation may serve as a warning sign for an adjustment of relative prices and the likelihood of a slowdown of the aggregate growth rate of the economy. According to Rodrick (2008) real exchange rate changes also indicate production and consumption taking place between domestic and foreign goods, and a misalignment of the real exchange rate can

be used as a method of influencing the actual state of the economy. Countries may try to maintain undervalued economies in an attempt to stimulate economic growth by increasing exports, capital flow and depreciating national currency. This has played a central role in the successful development of China (Coudert et al., 2007).

2.2 Real Equilibrium Exchange Rate: Theory and Methods and Impact of Misalignment on Economic Growth

Purchasing Power Parity is mostly used to determine long-term nominal interest rate, it is given by domestic price level divided by foreign price level. The conjecture forming bases of purchasing power parity is that the law of One Price assumed to be true for every good in the price basket((Rajan and Siregar 2002). The Natural Rate of Exchange rate this approach separate the medium-run and the long-run equilibrium real exchange rate. The medium-run is defined as a period where internal and external balance is achieved, the medium-run is obtained by using current capital stock with foreign debt values whilst the long-term equilibrium is constructed through sustainability of capital stock with foreign debt at a steady state(Egert 2004). The Natural Rate of Exchange rate model does not need a stable equilibrium real exchange rate, and it depends on real economic fundamentals prevailing in the economy(Rajan and Siregar 2002). Behavioural Real Exchange rate is a general approach used to model equilibrium exchange rate, it differentiate between exchange rate, economic fundamentals and short-run variables. Through calculating actual or current misalignment then set short-run variables to zero then substitute into the estimated relationship, misalignment will be the difference between the fitted values and actual values of the real exchange rate(Egert et al 2005). Fundamentals Equilibrium Exchange Rate is concerned with sustainable external equilibrium external account-based equilibrium real exchange rate. Fundamental equilibrium exchange rate is effective exchange rate that ensures internal and external balance of a country or countries or more simultaneously.

Kamar et al., (2007) explain that it is not easy to set, or determine the most suitable exchange rate, or to keep the exchange rate at the correct level. Some countries seeking to enter into a monetary union, such as the Eurozone, are required to keep their exchange rate floating at a predetermined parity, with approximately 15 percent deviation at most, for a period of two years (Magyari, 2008). Conversion rates to the Euro system determine the extent of the social and financial impact of the Euro changeover, as the rate of changeover must be similar to the rate at which the economy is moving, or the country's economic performance can be effected.

Faced with these situations, it is important to determine a correlation between the growth of the economy and the real exchange rate misalignment (Mastrobuoni, 2004). Existing literature notes that establishing the wrong level of the real exchange rate may cause economic agents to read the economy incorrectly, and consequently create distortions and instability in the economy. Even though this might be true, devaluation of the currency may promote economic growth by increasing exports, which stimulate economic productivity (Bayoumi and Mauro, 2001).

The literature on the impact of real exchange rate misalignment on economic growth uses panel data analysis, although depending on the interests of the researcher, the dependent variables could be real exchange rate misalignment, or economic growth (Rodrick, 2008). The performance of the indicators included in the studies is considered, as many of these studies use economic growth calculated through the growth rate of the gross domestic product, while in other studies they apply components of the gross domestic product (MacDonald and Vieira, 2010). In the literature many drawbacks have been noticed concerning the real exchange rate misalignment and some research uses purchasing power parity and the approach of the general equilibrium. With these two approaches important drawbacks are noted: for example the purchasing power parity approach cannot be empirically tested because there is not enough data available to permit analysis; and the general equilibrium approach is difficult to use because of specific structural issues in developing countries. Because the approach assumes it is modelling the whole world it is easy to apply (Magyari, 2008; Coudert and Couharde, 2009).

REER equilibrium misalignment estimated for SADC economies reveal continuous overvaluation (Zerihun, 2014). According to Magyari, (2008) there are two studies where real exchange rate misalignment is estimated by a single equation approach. Studies have found the correlation between real exchange rate misalignment and economic growth to be negative, which provides empirical proof that devaluation of a currency can enhance the performance of the economy and economic growth would be higher if new and different channels of transmission were to be found. One of the channels that has been investigated in the literature, is through investment by the rest of the world, as stimulating the accumulation of capital. Alternatively, according to Rodrick (2008) when the real exchange rate moves away from equilibrium it could have an impact on the trade of goods and the competitiveness of the economy with respect to the rest of the world.

Kamar and Naceur, (2007) used PMG estimation and found that money supply, budget deficit, government consumption and degree with which countries are involved in international trade had similar effects on the real exchange rate. In addition they determined that if there is misalignment in the short-run, the long run will be characterised by convergence. If there is interdependence between countries it is not appropriate to estimate PMG (Coleman, 2008); this is also shown by Pesaran et al., (1999). Real exchange rates consist of an important characteristic which serves as an appropriate measure for intensity of a country's international competitiveness (Edward, 1988), while Rodrik (2008) argues either a decline in real exchange rate is an indication that there might have been a rise in domestic costs of production of goods and services.

Real exchange rate misalignment is a function of factors such as openness, ratio of investment to GDP, terms of trade and when there is a rise in explanatory variables it leads to an appreciation of real exchange rates (Eita and Jordaan, 2013). Kamar and Naceur's (2007) theory shows that the real exchange rate can be influenced by a change of variables, such as monetary policy, government expenditure, terms of trade, degree of openness and capital flows. The real exchange rate has central importance to economic activity since it influences, and it is also influenced by other policies. Achieving policy coordination and harmonization is necessary for successfully establishing a common currency.

Real exchange rate misalignment is more likely to be better explained by panel data, and in situations where there is relatively minor differences between countries on actual real exchange rate and estimated real exchange rate. Misalignment implies that there is a lack of similarity between economies in the explanatory variables of real exchange rates (Dunaway et al., 2006). Empirical work done by MacDonald and Vieira (2010) using a GMM model produces positive estimates for all explanatory variables for real exchange rate misalignment, which means that real GDP growth for their study was influenced by real exchange rate depreciation, and real exchange rate appreciation is found to negatively influence real GDP growth rate. Bleaney and Greenaway (2001) suggest that real exchange rate misalignment has a negative impact on investment; also deviations in terms of trade impact negatively on the economic growth rate. Yotopoulos and Sawada (2006) used panel data when determining that real exchange rate misalignment from purchasing power parity deviations are unique from one country to the next.

In literature, the analysis of the exchange rate is done in two parts; undervaluation and overvaluation indicators, for which panel data is used when estimating. When observations are fitted on real GDP growth, it is established that the effect is negative and statistically significant on growth. However, the relationship between economic growth and real exchange rate undervaluation is not statistically significant, but the harmful influence of volatility of exchange rate misalignment is noticeable on economic growth (Rodrick, 2008). The literature measures real exchange rate misalignment as a deviation from equilibrium real exchange rate (MacDonald and Vieira, 2010). The literature also identifies three different ways that can be used to compare exchange rate misalignment. The first is a PPP-based measure of misalignment, which uses deviations from equilibrium real exchange rate with respect to real exchange rate determined in some year as equilibrium (Magyari, 2008). One disadvantage of using this approach is that purchasing power parity only takes into account changes in the exchange rate resulting from nominal variables, and ignores fluctuations of the exchange rate attributed to real factors.

Secondly, some scholars calculate exchange rate misalignment through differences between the black market and official exchange rates Magyari, (2008). Where a black market premium is used as proxy, it can better explain the extent to which foreign exchange controls may not be explaining real exchange rate misalignment during the time when the rest of the world is moving towards financial integration (Terra and Valladares, 2003). In addition, other studies have found a degree of exchange rate misalignment by the black market premium, and in developing countries in the 1970s and 1980s. Lastly there is a model based on the measure of real exchange rate misalignment; a theoretical equilibrium path in which misalignment is compared by deviations of actual real exchange rates (Aguirre and Calderon, 2005).

The model based measure of real exchange rate misalignment is based on the calculation of equilibrium exchange rate. Establishing real exchange rate equilibrium helps in determining simultaneously internal and external equilibrium. When looking at the literature most empirical work that has been done can be classified in a single equation model, and the general equilibrium simulation method. Both these measures treat real exchange rate as a relative price of traded goods and non-traded, by which it achieves internal and external equilibrium simultaneously (Magyari, 2008). In most instances a single equation method is derived as a reduced form for the equilibrium real exchange rate, basing it on a strong theoretical background (MacDonald and Vieira, 2010).

The application of a theoretical framework to derive real exchange rate long-run relationships in the existing literature links the real exchange rate to economic fundamentals, such as terms of trade, trade policy, productivity differentials, net foreign assets among others (Rodrick, 2008). In such situations misalignment happens when there are persistent deviations on real exchange rates from the equilibrium, as well as other factors caused by inadequate macroeconomic traded exchange rate policies (Terra and Valladares, 2003). Aguirre and Calderon, (2005) used a single equation approach to calculate real exchange misalignment. Initially this was done by estimating a long-run real exchange rate equation using historical data time series, and through panel data techniques. To estimate long-run values of the real exchange rate fundamentals, one can use different kinds of trend-cycle decomposition techniques or one can use a band-pass filter (King and Rebelo, 1993). Magyari (2008) used both HP-filters and band-pass filters.

Atingi-Ego and Sebudde (2004) used the Hodrick-Priscort filter to estimate equilibrium rate exchange rate, which represents a relative stability of existence of real exchange equilibrium. King and Rebelo (1993) explain that the Hodrick-Priscort filter is used to determine the implicit model that can be filtered to be optimal through minimizing mean square errors. Ravn and Uhlig (2002) suggest that when conducting a cross country analysis, the Hodrick-Priscort filter parameter should be adjusted according to the forth power of the frequency of observation. Scholars using the Hodrick-Priscort filter are more likely to have different estimations from researchers using differenced data (Cogley and Nason, 1995).

If a researcher is using band-pass filters to estimate equilibrium real exchange rate, estimated coefficients are multiplied with permanent values of the fundamentals. With the other approach used to compute real exchange rate misalignment, we find a difference between the actual and real exchange rate equilibrium (Rodrick, 2008). Alternatively, general equilibrium simulation models can be used to evaluate the behaviour of the real exchange rate. This method is different because the equilibrium real exchange rate meets both internal and external equilibrium conditions. The short fall of this approach is that it ignores important components such as the stock of demand for net foreign assets, but most simulations of this model are based on flow conditions (Magyari, 2008).

With links between real exchange misalignment and economic growth, some literature suggests that real exchange rate misalignment results in a negative effect on the allocation of resources, and therefore economic growth (Bayoumi and Mauro, 2001). However, some

researchers hold a premise that undervaluation could be an indication of competitive devaluation that can push exchange rates to a level that encourages an increase in exports (Magyari, 2008). Devaluation positively impacts on the economy because the economy grows through various new sources and adopts new technologies (Terra and Valladares, 2003). Agosin et al. (2012) suggest that there is a connection associated with real exchange rate and an increase with aggregate saving and investment, coupled with a decrease in unemployment, and real exchange rate depreciation, to stimulate economic growth.

Chapter 3

3.0. Methodology

The variables used in this study were obtained from the World Bank-Development Indicators data base. For the exchange rate misalignment section, the dependent variable is real exchange rate (REER). The independent variables are given by government consumption (GCON), budget balance (BUDG), degree of openness (OPEN), gross domestic product per capita (GDPPCAP), stock of reserves at the end of the year (RESY). All the variables are in log form except budget balance.

3.1. Unit Root testing

3.1.1 Testing for cross-sectional dependence

It is important to test for cross-sectional dependence, to confirm that it is appropriate to assume cross-sectional independence by Levein, Lin and Chu (LLC), Im-Pesaran-Shin (IPS), Fisher type unit root tests and by the Pedroni cointegration test. If dependence is found, a different unit root test and cointegrating test that allows for dependence must be used to obtain accurate estimates and that would mean we will have to estimate exchange rate misalignment using mean group (MG) instead of PMG. Let us consider a basic panel data regression

$$y_{RER_{it}} = \gamma_i + \rho' k_{it} + u_{it}, \quad i = 1, \dots, N \text{ and } t = 1, \dots, T \quad (1)$$

where k_{it} denote the vectors regressors $M \times 1$, ρ represents of parameter to be estimated which takes dimension $M \times 1$ and where γ_i denote a characteristic that is fixed across time. u_{it} is assumed to be independent and identically identified (i.i.d) across cross-sectional units and over-time period, where the assumption is the null hypothesis u_{it} . The alternative hypothesis is u_{it} may be correlated across cross-sectional units, however, the assumption of not autocorrelation remains.

$$H_0: \beta_{ij} = \beta_{ji} = \text{cor}(u_{it}, u_{jt}) = 0 \text{ where } i \neq j,$$

$$H_0: \beta_{ij} = \beta_{ji} \neq 0 \text{ where } i \neq j$$

if β_{ij} is generated by correlation coefficient of the distribution and is produced by

$$\beta_{ij} = \beta_{ji} = \frac{\sum_{t=1}^T u_{it} u_{jt}}{(\sum_{t=1}^T u_{it}^2)^{1/2} (\sum_{t=1}^T u_{jt}^2)^{1/2}}$$

$N \rightarrow \infty$ the resulting number of possible combinations of (u_{it}, u_{jt})

If we let N be fixed as $T \rightarrow \infty$, to test for Pesaran's CD test we have to use the Lagrange Multiplier (LM) statistic proposed by Breusch and Pagan (1980) which is also used by Hoyos and Sarafidis (2005) and Pesaran (2012). The statistic is given by

$$CD = T \sum_{i=1}^{N-1} \sum_{j=i+1}^N \hat{\beta}_{ij}^2$$

where $\hat{\beta}_{ij}^2$ represents the sample estimate of the pair-wise correlation of the residuals.

$$\beta_{ij} = \beta_{ji} = \frac{\sum_{t=1}^T u_{it} u_{jt}}{(\sum_{t=1}^T u_{it}^2)^{1/2} (\sum_{t=1}^T u_{jt}^2)^{1/2}}$$

where \hat{u}_{ij} is estimated of u_{it} , it is estimated by OLS using the regression

$$u_{it} = y_{RER_{it}} - \hat{\rho}' k_{it}$$

The LM follows a chi-square distribution and it is asymptotically distributed. It consists of degrees of freedom that are described by $N(N-1)/2$ with the application of the null hypothesis of interest. The LM statistic is not correctly centred when T is finite while N is larger and the bias gets worse as $N \rightarrow \infty$ with finite T , this argument is also affirmed by Baltagi et al (2007). The LM test is valid when N is small and T is sufficiently large. Under the null and the LM following a chi-square distribution $H_0, T\hat{\beta}_{ij}^2 \overset{a}{\sim} \chi_1^2$ with $\hat{\beta}_{ij}^2, i = 1, \dots, N-1$ $j = i+1, 2, \dots, N$, are asymptotically independent, therefore the scale version CD_{lm} is appropriate to use to test the hypothesis of cross dependence even in situations that consist of large N and large T .

$$CD_{lm} = \sqrt{\frac{1}{N(N-1)}} \sum_{i=1}^{N-1} \sum_{j=i+1}^N (T\hat{\beta}_{ij}^2 - 1)$$

Under the scaled version of the CD_{lm} with the null H_0 with $T \rightarrow \infty$ while at the same time $N \rightarrow \infty$, $CD_{lm} \xrightarrow{d} N(0,1)$, CD_{lm} is not correctly centered at zero. To correct this distortion we use the test proposed by Pesaran (2004)

$$CD = \sqrt{\frac{2T}{N(N-1)}} \left(\sum_{i=1}^{N-1} \sum_{j=i+1}^N T \hat{\beta}_{ij}^2 \right) \quad (2)$$

In the model above under the null hypothesis of no cross-sectional dependence $CD_{lm} \xrightarrow{d} N(0,1)$ even for $N \rightarrow \infty$ and T sufficiently large, this statistic is unlike the LM statistic CD has a mean of zero for fixed values of T and where there is a large panel size, including heterogeneous models non-stationary and dynamic models. When testing for cross-sectional dependence we use critical values on the paper published by Pesaran (2004).

Table 1 presents the result of the cross-section dependence test using the Pesaran (2004) CD test. Openness and stock of reserves shows independence while the remaining variables indicate that dependence exists between the cross sections. Therefore, it is necessary to take into account the existence of cross-sectional dependence while testing for unit root and cointegration.

Table 1 Cross-section Pesaran (2004) Dependence test

CD 2004 Pesaran	
Variables	Statistic
flnreer_172	7.07***
Bugd	3.64***
lngdppcap	12.77***
Lngexp	5.39***
Lnopen	-0.48
Lnresy	1.37

This is a test for cross-sectional dependence. Null hypothesis is cross sectional independence. *, **, *** denote significance at 10%, 5% and 1%.

3.2.1 Levin, Lin and Chu Test (LLC)

The LLC test states that each unit root has less power against alternative hypotheses, which have consistently higher deviations from equilibrium. This test is more powerful than testing for individual unit root test for each cross-section (Hlouskova and Wagner, 2006).

The LLC test was first proposed by Levin and Lin (1992, 1993) and later by Levin, Lin and Chu (2002). If we observe variables in N countries and for T periods, and initially we consider a model with individual fixed effect and no time trends, the model consists of a lagged dependent variable which is homogenous by restriction for all individuals of the panel

$$\Delta y_{i,t} = \alpha_i + \rho y_{i,t-1} + \sum_{z=1}^{p_i} \beta_{i,z} \Delta y_{i,t-z} + \varepsilon_{i,t} \quad (3)$$

Where $i = 1, \dots, N$ and where $t = 1, \dots, T$. The errors $\varepsilon_{i,t}$ *i.i.d.* $(0, \sigma_{\varepsilon_i}^2)$ by assumption are independent across all cross-sectional units in the sample. In specifying this model the interest is on testing the null hypotheses $H_0: \rho = 0$ against the alternative hypotheses $H_1: \rho = \rho_i < 0$ for all $i = 1, \dots, N$ applying the auxiliary assumption about the individual effect $\alpha_i = 0$ for all $i = 1, \dots, N$ under the null hypotheses $H_0: \rho = 0$. The alternative hypothesis is restrictive because it implies homogenous autoregressive parameters across all panels. In cases where we use the LLC to test for the convergence hypotheses in growth models, the alternative restricts every country to converge at the same rate. Although it is important to understand that using pooled estimator $\hat{\rho}$, even when DGP is not identical, does not mean that the unit root test is inconsistent. (Baltagi, 2008).

To illustrate this point, let us consider a sample linear model $y_i = x_i \beta_i + \varepsilon_i$ suppose β_i is equal to 0 in one of the samples and it is equal to 1 in the other half of the sample, assuming that we want to test for the null hypotheses $\beta_i = 0$ for all the units. This test is possible in the context of pooled estimate $\hat{\beta}$ on the entire sample (Baltagi, 2008). Hurlin and Mignon (2006) state that the pooled OLS would produce estimates that converge to 0.5 and the standard error would converge to zero, which means the null hypothesis will be rejected. Normally it is likely to obtain a more powerful test by splitting a sample into two separate parts and then conducting a test with the null hypothesis in both parts (Hlouskova and Wagner, 2006). Therefore it is important to split issues of estimating the value of the autoregressive parameter to be able to estimate the rate at which convergence is taking place,.

The null hypothesis is that all panels contain unit root against the alternative which is all panels contain unit root. The maintain hypothesis is that

$$\Delta \rho y_{i,t-1} + \sum_{L=1}^{p_i} \theta_{iL} \Delta y_{it-L} + \alpha_{mi} d_{mt} + \varepsilon_{it}, \quad m = 1, 2, 3 \quad (4)$$

Where d_{mt} is denoting the vector of a deterministic trend variables and α_{mi} denotes the corresponding vector of coefficient of the model $m = 1, 2, 3$. In a more specific definition $d_{1t} = \{\text{empty set}\}$, $d_{2t} = \{1\}$ and $d_{3t} = \{1, t\}$. The lag order p_i is not known. Baltagi (2008) argues that LLC suggests a three step approach to be followed when implementing the test. The first step in running the LLC as we need to perform a separate augmented Dicky-Fuller regression for all cross-sections individually (that is equation 4) the lag order p_i is allowed to differ across all individual cross-sections. Given the size of T , we choose the largest number of lag order p_{max} and take the estimated t-statistic of $\hat{\theta}_{iL}$ to evaluate whether a smaller lag order is accepted to a larger order, (t-statistics are normally distributed with a mean of zero and a variance equalling one) the null hypotheses ($\theta_{iL} = 0$), in both situations when $\rho_i = 0$ and $\rho_i < 0$.

Once the adequate lag order is determined, two auxiliary regressions are regressed to obtain orthogonalized residuals:

$\Delta y_{i,t}$ on $\Delta y_{i,t-L}$ ($L = 1, \dots, p_i$) and d_{mt} to get residuals \hat{e}_{it}

$\Delta y_{i,t-1}$ on $\Delta y_{i,t-L}$ ($L = 1, \dots, p_i$) and d_{mt} to get residuals \hat{v}_{it-1}

The residuals are standardized to control for differences of variances across i , $\hat{e}_{it} = \hat{e}_{it} / \hat{\sigma}_{\varepsilon i}$ and $\hat{v}_{it-1} = \hat{v}_{it-1} / \hat{\sigma}_{\varepsilon i}$. $\hat{\sigma}_{\varepsilon i}$ denotes standard error from each of the ADF regression, for $i = 1, \dots, N$. The next step is to estimate the short-run and the long run ratio of standard deviations, where the null hypothesis of a unit root for the long-run can be estimated in the following format

$$\hat{\sigma}_{y_i}^2 = \frac{1}{T-1} \sum_{t=2}^T \Delta y_{it}^2 + 2 \sum_{L=1}^{\bar{K}} w_{\bar{K}L} \left[\frac{1}{T-1} \sum_{t=2+L}^T \Delta y_{it} \Delta y_{i,t-L} \right] \quad (5)$$

where the truncated lag is denoted by \bar{K} (in this study there is no truncation because data was available for all years) which might be dependent on data. \bar{K} is estimate in a way that

guarantees $\hat{\sigma}_{yi}^2$ is consistent. In each of the cross-section i , $w_{kl} = 1 - (\frac{L}{K+1})$, where the ratio showing the long-run standard deviation to the innovation standard deviation is given by $\hat{s}_{it} = \hat{\sigma}_{yt}/\hat{\sigma}_{\varepsilon i}$. The estimation of the average of standard deviation is given by $\hat{S}_N = \frac{1}{N} \sum_{i=1}^N \hat{s}_i$. Then to conduct a panel test statistic requires that we run the pooled regression.

$$\tilde{e}_{it} = \rho \tilde{v}_{i,t-1} + \tilde{\varepsilon}_{it} \quad (6)$$

Based on a given $N\tilde{T}$ observations where $\tilde{T} = T - \bar{p} - 1$. \tilde{T} denotes the number the average number of each individual in the panel with $\bar{p} = \sum_{i=1}^N p_i / N$, where \bar{p} is defined as the average lag order of each individual ADF regression. The common t-statistic for the null hypotheses for, $H_0: \rho = 0$ is defined by $t_\rho = \frac{\hat{\rho}}{\hat{\sigma}(\hat{\rho})}$

where;

$$\hat{\rho} = \frac{\sum_{i=1}^N \sum_{t=2+p_1}^T \tilde{v}_{i,t-1} \tilde{e}_{it}}{\sum_{i=1}^N \sum_{t=2+p_1}^T \tilde{v}_{i,t-1}^2}$$

$$\hat{\sigma}(\hat{\rho}) = \hat{\sigma}_{\tilde{\varepsilon}} / \left[\sum_{i=1}^N \sum_{t=2+p_1}^T \tilde{v}_{i,t-1}^2 \right]^{1/2}$$

and

$$\hat{\sigma}_{\tilde{\varepsilon}}^2 = \frac{1}{N\tilde{T}} \sum_{i=1}^N \sum_{t=2+p_1}^T (\tilde{e}_{it} - \hat{\rho} \tilde{v}_{i,t-1})^2$$

This equation is to show how the variance is estimated, the adjusted t-statistic is defined by,

$$t_\rho^* = \frac{t_\rho - N\tilde{T}\tilde{S}_N \tilde{\sigma}_{\tilde{\varepsilon}}^{-2} \hat{\sigma}(\hat{\rho}) \mu_{m\tilde{T}}^*}{\sigma_{m\tilde{T}}^*} \quad (7)$$

The mean and standard deviation adjustments are denoted by $\mu_{m\tilde{T}}^*$ and $\sigma_{m\tilde{T}}^*$, t_ρ^* is said to be distributed asymptotically $N(0,1)$. The asymptotic condition requires $\sqrt{N_T}/T \rightarrow 0$, in this case N_T highlight the situation where the cross-sectional dimension N is assumed to be an arbitrary monotonically rising function of T (Hlouskova and Wagner, 2006). According to Baltagi (2008), this is applicable to micro panel data for a situation where the pace of the growth of T is allowed to be slower than the pace of growth of N_T . Other speeds of

divergence are sufficient, but not necessary such as $N_T/T \rightarrow 0$ and $N_T/T \rightarrow \text{constant}$. It is determined above that when using the LLC approach it is required that we specify the number of lags that we are going to make use of in each individual cross-section. ADF regression(p_i), and kernel choices are used in the computation of S_N .

We specify an exogenous variable that is going to be used in the test equation, where we can regress an exogeneous plus a trend, or exogeneous variable with no trend or no trend and no constant. When conducting panel unit root testing the acceptable size of N is between 10 and 250, and the moderate size for T is between 25 and 250. Baltagi (2008) remarks that it might be difficult to compute the standard panel procedures or it may not have enough power for a panel of this size. Although for a very large T , when testing for individual unit root time-series, the test will have enough power to apply for each individual cross-section, therefore for a very large N and small T the usual panel data procedures are preferred (Hurlin and Mignon, 2006). The use of a Monte Carlo simulation conducted by LLC shows that if data is normally distributed a better approximation of empirical distribution of the t-statistic occurs, even in a situation of relatively small sample size. Additionally the panel unit root test provides considerable increase in power compared to a separate unit root test for each of the cross-sections (Hoang and McNown, 2006).

The proposed LLC test like all other approaches is not without limitations. The validity of the LLC test depends on the independence assumption in the entire cross-section and the test is invalid if cross-section correlation is present. Additionally there is a restrictive assumption of all cross-sections do not or contain unit root. Hoang and McNown (2006) argues that with LLC if correlation is present in the cross-section, the test suffers from dramatic size distortion among contemporaneous cross-sectional error terms. Therefore it is important to control for cross-sectional dependence when running panel unit root testing of exchange rate.

3.2.2 Im, Pesaran and Shin Test (IPS)

We have determined from above that Liven, Lin and Chu is a restrictive test that requires ρ to be homogeneous across i . The LLC test is suggested to be most suitable for testing for convergence in growth among countries, but the disadvantage of this test is that the alternative is restrictive; it restrict countries to converge with the same rate (Baltagi, 2008). On the other hand the IPS permits heterogeneous coefficient of y_{it-1} thereby introduces an alternative testing approach that is based on the individual unit root test statistic. The IPS

argues that the average of the ADF test used when testing u_{it} contains autocorrelation (Canning and Pedroni, 2004), and for the model has different autocorrelation properties for all cross-sectional units,

$$\Delta \rho y_{i,t-1} + \sum_{L=1}^{p_i} \theta_{iL} \Delta y_{it-L} + \alpha_{mi} d_{mt} + \varepsilon_{it}, \quad m = 1, 2, 3 \quad (8)$$

As we have seen previously the null hypothesis is that all panels contain unit root, $H_0: \rho_i = 0$ for all cross-sectional units i , against the null hypothesis give way for some, but not all, of the individual series to have units roots, which is;

$$H_0: \begin{cases} \rho_i < 0 \text{ for } i = 1, 2, \dots, N_1 \\ \rho_i = 0 \text{ for } i = N_1 + 1, \dots, N \end{cases} \quad (9)$$

This test requires that the fraction of the individual time series that are stationary should not be zero, $\lim_{N \rightarrow \infty} (N_1/N) = \delta$ where $0 < \delta \leq 1$. This requirement is important because it ensures we obtain consistency when testing the panel for unit root test (Hoang and McNown, 2006). The t-bar statistic for IPS is defined as the average of the individual ADF statistic;

$$\bar{t} = \frac{1}{N} \sum_{i=1}^N t_{\rho_i} \quad (10)$$

where t_{ρ_i} denotes the individual t-statistic for conducting a null hypothesis test, $H_0: \rho_i = 0$ for all i in (9). In such situations the lag order is always zero ($p_i = 0$ for all i), the IPS offers simulated critical values for \bar{t} for a different number of cross-sections N and T number of time series where a Dickey-Fuller regression with a constant or constant and deterministic trends. The most common use is where the lag order p_i might not be zero for some of the cross-sections, the IPS indicates that an adequately standardized \bar{t} contains an asymptotic $N(0,1)$ distribution (Hlouskova and Wagner, 2006; Baltagi 2008). Supposed we start from a commonly known results time series with N fixed

$$t_i \Rightarrow \frac{\int_0^1 W_{iz} dW_{iz}}{\left[\int_0^1 W_{iz}^2 \right]^{1/2}} = t_{iT} \quad (11)$$

as $T \rightarrow \infty$, where $\int W(r)dr$ represents a Weiner integral that has augmented r suppressed in equation (11) (Billingsley, 1961). IPS operate under the assumption that t_{iT} i.i.d and consists of a finite mean and variance. Therefore,

$$\frac{\sqrt{N} \left(\frac{1}{N} \sum_{i=1}^N t_{iT} - \frac{1}{N} \sum_{i=1}^N E[t_{iT} | \rho_i = 0] \right)}{\sqrt{\frac{1}{N} \sum_{i=1}^N \text{var}[t_{iT} | \rho_i = 0]}} \Rightarrow N(0,1) \quad (12)$$

as $N \rightarrow \infty$, we apply the Lindeberg-Levy central theorem, so

$$t_{IPS} = \frac{\sqrt{N} \left(\bar{t} - \frac{1}{N} \sum_{i=1}^N E[t_{iT} | \rho_i = 0] \right)}{\sqrt{\frac{1}{N} \sum_{i=1}^N \text{var}[t_{iT} | \rho_i = 0]}} \Rightarrow N(0,1) \quad (13)$$

sequentially as $T \rightarrow \infty$ followed by $N \rightarrow \infty$, the values given by $E[t_{iT} | \rho_i = 0]$ and by $\text{var}[t_{iT} | \rho_i = 0]$ are produced by IPS through simulation for different values of T and p'_i 's. The Monte Carlo simulation reveals that when the lag order chosen is large enough for the ADF regression under consideration compared to a small sample performance of the t-bar test, IPS is reasonable satisfactory and generally better than the LLC test we have seen above (Hoang and McNown, 2006; Hlouskova and Wagner, 2006; Billingsley, 1961).

3.2.3 Fisher Type Unit Root Test

As mentioned previously the unit root test for panel data is based on a heterogeneous model and includes testing the significance of the result from N independent individual test. We have seen that the IPS test uses an average statistic, or we can use the alternative testing approach which is conducted by joining together the observed significant levels of the individual tests. This approach has a strong historical background and is based on the p-values, let us consider equation $\Delta y_t = \phi y_{t-1} + \varepsilon_t$ assumed to be heterogeneous model. The hypothesis is similar to the one we have above,

$H_0: \rho_i = 0$ for all $i = 1, \dots, N$ and the alternative hypothesis is $H_1: \rho_i < 0$ for $i = 1, \dots, N_1$ and $\rho_i = 0$ for $i = N_1 + 1, \dots, N$, with $0 < N_1 \leq N$. To illustrate the idea of the Fisher type test we consider the pure time series unit root test statistic, for continuous statistics, the corresponding p-values are denoted by p_i which are uniform variables (0,1) (Maddala and

Wu 1999). Since we make an important assumption that there is cross-sectional independence, defined as

$$P_{MW} = -2 \sum_{i=1}^N \log p_i \quad (14)$$

The test follows a chi-square distribution comprising of $2N$ degrees of freedom as $T \rightarrow \infty$ for a given value of N . This test is simple and provides a robust statistical alternative. Sample size and lag length render this approach extremely favourable to use for panel unit root testing. When we have a large N sample, Hurlin and Mignon (2007) suggest this standardised approach is used.

$$\begin{aligned} Z_{MN} &= \frac{\sqrt{N}\{N^{-1}P_{MN} - E[-2 \log p_i]\}}{\sqrt{\text{Var}[-2 \log p_i]}} \\ &= -\frac{\sum_{i=1}^N \log p_i + N}{\sqrt{N}} \end{aligned} \quad (15)$$

They argue that this test statistic is consistent with the standardised cross-sectional average of individual p-values, hence the Lendeborg-Levy theorem is sufficient to show convergence to the standard normal distribution in the case of the unit root hypothesis (Hoang and McNown, 2006; Hlouskova and Wagner, 2006).

Billingsley (1961) states that the central limit theorem proposed by Lendeborg and Levy state that if $\{u_1, u_2, \dots\}$ is i.i.d order of white noise variables that has finite second moments, which makes the distribution of $n^{-1/2} \sum_{k=1}^n u_k$ follow the normal distribution with a mean of zero and the variance $E\{u_1^2\}$, with the assumption that $E\{u_1\} = 0$. The assumption of independence would be weakened under these conditions.

If we let $\{u_1, u_2, \dots\}$ to be stationary, and allow a positive recurrent aperiodic case of the stochastic process such that it produces $E\{u_1^2\}$ which is finite and $E\{u_n | u_1, \dots, u_{n-1}\} = 0$ (16) with a probability of one, and then $n^{-1/2} \sum_{k=1}^n u_k$ is distributed in a way that approaches the normal distribution with a mean equalling zero and variance $E\{u_1^2\}$ (Hlouskova and Wagner, 2006).

Condition (16) satisfies the requirement that the partial sums $\sum_{k=1}^n u_k$ create a martingale. To make this theorem hold assume that Ω denotes the Cartesian product representing the order of

copies of real line, represented by the integers $n = 0, \pm 1, \pm 2, \dots$. The coordinate variable is indexed by u_n , the Borel field is denoted by β and the P is the likelihood that estimate the finite-dimensional distributions introduced by initial processes (Hoang and McNown, 2006; Hlouskova and Wagner, 2006). If we suppose that f_n is the Borel field given by $\{u_n, u_{n-1}, u_{n-2}, \dots\}$, through condition (16)

$E\{u_n | f_{n-1}\} = 0$, (17) with a probability of one for integers $n = 0, \pm 1, \dots$. If $\sigma_n^2 = E\{u_n^2 | f_{n-1}\}$ and where $\sigma^2 = E\{\sigma_n^2\} = E\{u_n^2\}$, where T is the translation operator (in times series we refer to this as the lag operator), therefore it becomes simple to show that, $\sigma_n^2 = T^n \sigma_0^2$. Since T is a random process that time averages of a single sequence of events therefore it is consistent with the ergodic theorem. Condition (18) $\lim_{n \rightarrow \infty} n^{-1} \sum_{k=1}^n \sigma_k^2 = \sigma^2$ assumed to have a probability of one. If $s_n^2 = \sigma_1^2 + \dots + \sigma_n^2$, let $m_t = \min\{n: s_n^2 \geq t\}$ for all values of $t > 0$, where c_t is the value such that it is $0 < c_t \leq 1$ and $s_{m_t-1}^2 + c_t^2 \sigma_{m_t}^2 = t$, in addition we allow $z_t = u_t + \dots + u_{m_t-1} + c_t u_{m_t}$. Because of condition (19) it is determinable that $\sum_k \sigma_k^2$ moves away (i.e. does not) converge with a probability equalling unity, making m_t and other variables to be adequately defined (Billingsley, 1961). A common premise of renewal type used in equation (19) indicates that

$$\frac{\lim_{t \rightarrow \infty} m_t}{t} = \sigma^{-2} \quad (20)$$

Therefore

$$P \left\{ n^{-1/2} \left| \sum_{k=1}^n u_k - z_n \sigma^2 \right| > \epsilon \right\} \leq (1 + 8\sigma^2) \epsilon$$

If we are able to obtain $u_n = f(z_n)$, where z_n is defines the Markov process satisfying $\{u_n, u_{n-1}, u_{n-2}, \dots\}$, additionally when $E\{f(z_n)^2\}$ is finite in the case when z_1 contains a stationary distribution. This argument can be supported by showing that $n^{-1/2} \sum_{k=1}^n u_k$ is asymptotically normal whether or not the distribution of z_1 is stationary

Table2 below presents the results for unit root testing using IPS, LLC and Fisher tests type. The IPS shows that all variables are insignificant at all levels except for the log of real exchange rate which is significant at 1% level of significant when there is no trend or constant. When a constant and trend is included all variables are insignificant, therefore these variables contain unit root.

When using the LLC with trend and no intercept, log of real interest rate is significant, and the government budget balance is significant at 5%; net capital flow and net foreign assets are significant at 10%. When a trend is included, all variables are insignificant. We therefore conclude that these variables have unit root.

For the Fisher unit root test, openness is said to be stationary while using the intercept only. However, when we consider the intercept and trend option, all the variables contain a unit root.

Table 2 Unit root test IPS, LLC and Fisher type test

Variables	IPS		LLC		Fisher	
	Intercept	Intercept+Trend	Intercept	Intercept+Trend	Intercept	Intercept+Trend
lnreer_172	-1.3240***	-1.359	-1.9093**	-2.7403***	14.0282	13.2558
bugd	-1.2076	-1.1224	-2.2446**	-2.7846***	13.1414	11.9593
lngdppcap	3.6528	0.5537	1.6892	0.0123	2.4472	7.9516
lngexp	0.497	1.5429	0.0821	0.2765	8.1444	2.1515
lnm2	0.3356	0.3851	-0.7912	-1.5591*	9.1912	5.8634
lnopen	-0.7696	-0.4823	-0.9765	-2.8189*	16.7146***	8.948
lnresy	0	0	-0.9044	-3.0128*	9.9623	21.8587
LNTOT	2.0742	-0.4407	0.7365	-2.7605*	1.7361	9.5423
NFA	-0.7295	-0.519	-1.4588*	-1.6526**	11.1047	8.0682
NKF	-0.7086	-0.7652	-1.3028*	-1.2479	11.0446	9.8533

Note: LLC, IPS and FISHER statistics correspond to a test of the null hypothesis that all the panels contain unit roots against alternative. *, **, *** denote significance at 10%, 5% and 1%. 2 lags were used to compute the test statistic.

3.3 Panel Unit Root Testing Allowing For Cross-Sectional Dependence

To deal with the problem of cross-sectional dependence the test explained above considers a one factor model, with a heterogeneous loading factor for error terms. The test estimates the ADF model consisting of cross-sectional average of lagged and they are integrated of order one as individual series (Breitung, and Das 2005). If there is no serial correlation of residuals the regression is described in the following way

$$\Delta q_{i,t} = \beta_i + \theta_i q_{i,t} + \sigma_i \bar{q}_{t-1} + d_i \Delta \bar{q}_t + \varepsilon_{i,t}$$

where we define $\bar{q}_{t-1} = \frac{1}{N} \sum_{i=1}^N q_{i,t-1}$ and $\Delta \bar{q}_t = \frac{1}{N} \sum_{i=1}^N q_{i,t}$, let $t_i(N, T)$ denote the test statistic obtained from estimating σ_i by using OLS. Unit root testing is now based on cross-sectional units root test ADF statistics (CADF) (Bailey et al, 2014). Pesaran (2003) contends that under extreme cases some values of T may be truncated to avoid having a small value of T in a sample. It follows that we are now able to construct a suitable version of the IPS t-bar test that is able to account for dependence. The modified t-bar is based on a CADF average of individual statistics (see Hurlin and Mignon, 2006). The cross-sectionally augmented IPS is defined as

$$CIPS = \frac{1}{N} \sum_{i=1}^N t_i(N, T)$$

Under extreme situations a truncated CADF statistic is described in the following way;

$$t_i^*(N, T) = \begin{cases} F_1 & \text{si } t_i(N, T) \leq F_1 \\ t_i(N, T) & \text{si } F_1 < t_i(N, T) < F_2 \\ F_2 & \text{si } t_i(N, T) \geq F_2 \end{cases}$$

where F_1 and F_2 represent intercepts that are fixed to increase the likelihood that $t_i(N, T)$ associated with $[F_1, F_2]$ is close to unity. For this paper we are not going to truncate the time because the sample consists of the data starting from 1995-2012, while there are only 5 cross-sectional units. However it is worth noting that truncated data is characterised by a similar asymptotic null distribution that does not depend on the loading factor (Breitung and Pesaran 2008). We use simulated critical values for CIPS with two lags chosen by the SIC.

Table 3 shows the results for panel unit root testing while allowing for cross-sectional dependence using the Pesaran (2007) CIPS test. Two lags which were chosen by the SIC.

Under the null hypothesis, the series contains a unit root and we can see that when we include an intercept with no trend, log of government expenditure, lag 2 is significant at 1% level of significance. Therefore we reject the null hypothesis that the series contains a unit root. The CIPS test is generating insignificant results for all variables for all lags Bugd, lngdppcap, Lnopen and Lnresy are insignificant at all conventional levels of significance, and the null hypothesis cannot be rejected. lnreer_172 is insignificant at all levels of significance it can be concluded that the series contains units roots.

If we include a constant and a trend, Lnresy is significant at 1% level of significance 2 lags the null hypothesis can be rejected and we conclude the series contains unit root. However, the variables lnreer_172, Bugd, lngdppcap, lngexp and lnopen are insignificant at all conventional levels of significance, and the null hypothesis of the series contains unit root. The results have not changed in any considerable form from those produced by the unit root test that are unable to account for dependence between countries, but it is clear they are slightly different with reference to the log of real exchange rate the under the IPS and LLC, where they were significant. When using CIPS we find that the variable is insignificant; LLC, IPS and Fisher type are not able to detect structural breaks, and if there has been a structural break the power of these test is reduced. Hong and McNown, (2006) affirm that the power of LLC and IPS is low if there is a structural break.

Table 3 Panel Unit Root test (CIPS) Pesaran (2007)

Pesaran (2007) Panel Unit Root test (CIPS)			
Variables	Lags	Intercept	Intercept+Trend
lnreer_172	2	1.496	1.302
Bugd	2	1.935	-1.143
lngdppcap	2	0.395	1.67
lngexp	2	4.158***	3.472
lnopen	2	0.558	0.972
Lnresy	2	2.798	0.096***

This is a test that the null hypothesis of the series contains unit root. The CIPS test assumes that there is cross-sectional dependence in the form of a single unobserved common factor. .*, **, *** denotes significance at 10%, 5% and 1%. Only results on 2nd lag presented.

3.4 Pedroni Cointegration test

Let consider the most general form of a regression,

$$y_{it} = \alpha_i + \delta_i t + \beta_i X_{it} + e_{it} \quad (21)$$

in a case where there is a time series panel of observed y_{it} and X_{it} for cross-sectional units $i = 1, \dots, N$ in a certain period of time $t = 1, \dots, T$, where X_t is the column vector for each cross sectional unit i and β_i is also m -dimensional and is a row vector for each cross-sectional unit i . The variables y_{it} and X_{it} are nonstationary at level but it is assumed that they are integrated of order one indexed by $I(1)$, for each unit i of the panel, where the null hypothesis is there is no cointegration in the residuals e_{it} and will be $I(1)$ as well (Pedroni, 2001). The indexes α_i and δ_i give way to the possibility that cross-sectional units contain specific fixed effects and deterministic trends respectively. The slope of the coefficient of β_i is also allowed to change by individual, in order for the general cointegrating vector to be heterogeneous across member of the panel (Pedroni 2004).

With equation (24) the interest is in analysing the properties of the test for the null hypothesis H_0 : all of the individual panels are not cointegrated, and it is important to understand that the concerned data generating process (DGP) assumes that all individual panels are required to be uniformly cointegrated. Therefore the interpretation of the alternative hypothesis H_1 : is that all of the individual panels are cointegrated (Pedroni, 2004). The asymptotic equivalence does not hold in situations where regressors are endogenous therefore panels will no longer consist of homogenous coefficients, and it becomes important that we adjust to accommodate for the asymptotic bias that is initiated by the estimated regressors' effect. The properties of a test adjusted for a biased term under the null hypothesis of no cointegration, which is the result of estimated regressors, having influence in certain cases when there is endogeneity and where both slope estimates are limited to be homogeneous for all cross-sectional units (Pedroni, 2001).

This approach introduces a problem of interpreting the results of a null hypothesis test of no cointegration, more especially when the true data generating process (DGP) produces different slope coefficients across cross-sectional units. If we ignore differences in slope coefficients and impose a common slope, this permits estimated residuals for any cross-sectional unit with a different slope from the long-run regression correlation, to be non-stationary, even if there is cointegration in reality. Pedroni (2004) contends that in many

cases true slope coefficients are commonly heterogeneous across units of the panel, which makes the implication of adjusting slope coefficients to make them similar, not easily welcomed when testing for the null hypothesis of no cointegration.

Pedroni (1999; 2004) advocated the use of a test statistic that is derived from the stricter of residuals under the null of there is no cointegration that does not force slope coefficients of a regression to be the same for all units in the panel. The test statistic used on the null hypothesis of no cointegration in the general situation of the regression, are completely endogenous and they allow for differences in the slopes of coefficients for all units in the panel. Since we allow differences and cointegration to change across units of the panel, we can assume that the test is effectively pooling relevant information regarding the possibility of the presence of cointegrating relationships similar to those shown by the stationary characteristics of estimated residuals (Breitung, and Pesaran 2008).

To examine the properties of the distribution of the residual based test, let us define that DGP in terms of the partitioned vector $z'_{it} = (y_{it}, X'_{it})$ where the true process of z_{it} is defined by $z_{it} = z_{it-1} + \xi_{it}$, such that $\xi'_{it} \equiv (\xi^y_{it}, \xi^{x'}_{it})$ (Pedroni 2004). Assume for each cross-sectional unit i Invariance Principle, the standard functional central limit theorem is true for each member of the series as T approaches large values (Breitung, and Pesaran 2008). The DGP $\xi'_{it} \equiv (\xi^y_{it}, \xi^{x'}_{it})$ satisfies $1/\sqrt{T} \sum_{t=1}^{[Tr]} \xi_{it} \Rightarrow B_i(\Omega_i)$, for each cross-sectional unit as $T \rightarrow \infty$, in a case where \Rightarrow completes the condition of weak convergence and where $B_i(\Omega_i)$ is the vector Brownian motion containing an asymptotic covariance Ω_i such that the dimension is $m \times m$ lower diagonal block $\Omega_{22i} > 0$ and we further consider $B_i(\Omega_i)$ to be defined in the same dimensional space for all cross-sectional units i (Phillip and Moon, 1999). For this convergence to take place demands that the condition on the error process is relatively weak and includes the cluster of all stationary autoregressive moving average (ARMA) processes.

The asymptotic matrix $(m + 1) \times (m + 1)$ is a given by $\Omega_i \equiv \lim_{T \rightarrow \infty} E[T^{-1}(\sum_{t=1}^T \xi_{it})(\sum_{t=1}^T \xi'_{it})]$ and we can break it down as follows $\Omega_i \equiv \Omega_i^0 + \Gamma_i + \Gamma_i'$, where Ω_i^0 denotes the contemporaneous and Γ_i represents the dynamic covariance of ξ_{it} for a given cross-sectional unit i (Phillip and Moon, 1999; Pedroni 2004). The matrix is partitioned to ensure that it is consistent with the dimension of the vector $\xi'_{it} \equiv (\xi^y_{it}, \xi^{x'}_{it})$ in order the Ω_{22i} element portrays an m -dimensional matrix $m \times m$. The feedback between the regressors and the dependent variable is captured by the off-diagonal terms Ω_{21i} and x_{it} is not

allowed to be exogenous (Pedroni 2001). This is done to maintain consistency with existing literature. Since Ω_i is allowed to be heterogeneous across cross-sectional units, this is an indication that the panel are consumed in the asymptotic covariance matrix to be different. If we allow $\Omega_{22i} > 0$, we are disregarding the possibility of the situation of regressors being cointegrated with one another in a case where we have multiple regressors present (Kao, 1999).

With the assumption of cross-sectional independence (this condition for invariant principle is in reference to the time series demission), the individual cross-sectional units are assumed to be i.i.d. In a cross-section such that $E[\xi_{it}\xi'_{js}] = 0$ for all $s, t, i \neq j$. In simple terms the asymptotic long-run variance matrix of the size $N \times T$ panel is block diagonal with an i th block provided by asymptotic covariance for the member i giving a diagonal $(\Omega_1, \dots, \Omega_N)$ (Pedroni 2001). The error ξ_{it} is postulated to be given by a linear process $\xi_{it} = C_1(L)\eta_{it}$, in the case where Ω_i is described as $\Omega_i = C_i(1)C_i(1)'$ and η_{it} indexes the stochastic innovation, where the $C_1(L)$ denotes random coefficients that are independent from one another for both the i and t dimentions (Breitung, and Pesaran 2008). Therefore standard central limit theorem can be applied in the cross-sectional dimension containing different error terms in a relatively normal way. Suppose $C_1(L)$ is derived from an i.i.d distribution from dimension i which is independent from the innovation η_{it} , and because $\Omega_i > 0$ guarantees that between y_{it} and X_{it} there is no cointegrating relationship (Pedroni 2004).

Assumptions made above permit the use of the standard asymptotic convergence established from the time series dimension of each individual unit(Kao, 1999). Convergences hold for each individual unit $i = 1, \dots, N$ as T gets bigger.

If we allow $\{y_t\}_1^\infty$ to contain an h -dimension for a series generated by $y_t = Ax_t + u_t$ (22) where index A is an $n \times m$ coefficient matrix and where vector m process $\{x_t\}_0^\infty$ satisfies $x_t = x_{t-1} + v_t$ (23) where x_0 is considered to be any random variable, a constant is also included (Pedroni 2004). Then we extend condition (22) to become $y_t = \mu + Ax_t + u_t$ (1)' and $y_t = \mu + \theta t + Ax_t + u_t$ (24)'', where $w'_t = (u'_t, v'_t)$, then we need a small sum process $S_t = \sum_l^t w_j$ generated by the sequence $\{w_t\}_1^\infty$ which is consistent with the invariance principle, if $r \in [0,1]$ we describe $X_t(r) = T^{-1}S_{[Tr]}$, for this to make sense we need $X_t(r) \Rightarrow B(r)$ as $T \rightarrow \infty$, where \Rightarrow represents weak convergence for a corresponding

probability measure, and the indexes $B(r)' = (B_1^n(r), B_2^m(r)')$ is vector $(n + m)$ Brownian motion with a covariance matrix (Pedroni 2001; Kao, 1999).

$$\begin{aligned}\Omega &= \begin{bmatrix} \Omega_1 & \Omega'_{21} \\ \Omega_{21} & \Omega_2 \end{bmatrix} \begin{matrix} n \\ m \end{matrix} \\ &= \lim_{T \rightarrow \infty} T^{-1} E(S_t S_T') \\ &= \Sigma + \Lambda + \Lambda' \end{aligned} \tag{25}$$

where,

$$\Sigma = \begin{bmatrix} \Sigma_1 & \Sigma'_{21} \\ \Sigma_{21} & \Sigma_2 \end{bmatrix} \begin{matrix} n \\ m \end{matrix} = \lim_{T \rightarrow \infty} T^{-1} \Sigma_1^T \Sigma(w_t w_t') \tag{26}$$

and

$$\Lambda = \begin{bmatrix} \Lambda_1 & \Lambda'_{21} \\ \Lambda_{21} & \Lambda_2 \end{bmatrix} \begin{matrix} n \\ m \end{matrix} = \lim_{T \rightarrow \infty} T^{-1} \Sigma_{t=2}^T \Sigma_{j=1}^{t-1} \Sigma(w_t w_t') \tag{27}$$

Where $B_1(r)$ and $B_2(r)$ are two vectors of the Brownian motion, which in this respective order are dimension n and m with the covariance matrices Ω_1 and Ω_2 of which are assumed to be positive (Pedroni 2004). An invariant principle, such as the one show in (25) consists of a wide variety of sequences $[w_t]$ that are likely to be distributed differently and are weakly dependent (Kao, 1999). Condition (25) also applies to a large linear process and to other processes generated by all stationary investable ARMA models. The asymptotic theory regression depends on a weak convergence of particular sample covariance matrices to white noise innovation, integral of the form $\int_0^1 B dB'$

$$T^{-1} \sum_1^T S_{t-1} w_t' \Rightarrow \int_0^1 B dB' + \Lambda. \tag{25}'$$

We assume that condition (25) hold and that $\{w_t\}$ is stationary also ergodic consisting of fourth-order moments, additionally it follows that the time series $\{x_t\}$ is integrated of order one (Breitung, and Pesaran, 2008). Since we assume stationarity condition (25) and (26) are broken down to $\Sigma = E(w_1 w_1')$ and $\Lambda = \Sigma_2^\infty(w_1 w_1')$, if the series describing Λ converges absolutely it follows that $\{x_t\}$ contains a spectral density matrix $f_{ww}(\lambda)$ and this allows us to express $\Delta y_t = \Phi y_{t-1} + \varepsilon_t$ in the form,

$$\Omega = 2\pi f_{ww}(0).$$

Model (25) can be considered as a multivariate equation system with regressors x_t are generated by a general process of the first order, and x_t is strictly exogenous (Pedroni 2001). It follows that models described by condition (2) and condition (26) precisely imply that the time series $\{x_t\}$ and $\{y_t\}$ are integrated. Therefore we have shown that

$$T^{-2} \sum_{t=1}^T z_{it-1} z'_{it-1} \Rightarrow L'_i \int_0^1 Z_i(r) Z_i(r)' dr L_i, \quad (28)$$

$$T^{-2} \sum_{t=1}^T z_{it-1} \xi'_{it-1} \Rightarrow L'_i \int_0^1 Z_i(r) Z_i(r)' dr L_i + \Gamma_i, \quad (29)$$

exhibit weak convergence. Where $Z_i(r) \equiv (V_i(r), W_i(r)')'$ denote vector Brownian motion, such that $V_i(r)$ and $W_i(r)$ are independent for all units i , where that $W_i(r)$ is $m \times 1$ dimensional vector, Γ_i is the contemporaneous covariance vector as was defined before and L_i is a lower triangular decomposition of Ω_i ,

$$L_{11i} = (\Omega_{11i} - \Omega'_{21i} \Omega_{22i}^{-1} \Omega_{21i})^{1/2}, \quad L_{12i} = 0$$

$$L_{11i} = \Omega_{22i}^{-1} \Omega_{21i}, \quad L_{22i} = \Omega_{22i}^{1/2}$$

There is convergence and it results in equation (28) and (29), if we assume initialization of $z_{i0} = 0$ for all individual units i . Since we assume cross-sectional independence, we average cross-sectional sum of the Brownian motions' functions used to develop the panel statistics (Pedroni 2004). As $T \rightarrow \infty$ we obtain N , we get the sum statistic $P_{NT} = \sum_{i=1}^N \sum_{t=1}^T Y_{it}$ this sum is used in the construction of the panel statistic, so to simplify this we can write it as $P_{NT} = \sum_{i=1}^N S_{iT}$ where we let $S_{iT} = \sum_{t=1}^T Y_{it}$ (Pedroni 2001). If we allow R_i to represent the limit of standardized sum of S_{iT} as T gets large. To obtain the sequential limit we calculate R_i as T gets large, it follows that we calculate the limit of the standard sum of $\sum_{i=1}^N R_i$ as N increases. The application of the sequential limit theorem as $N \rightarrow \infty$ is optimal in determining the desired limit distribution for nonstationary double index theorem (Phillip and Moon, 1999), but Pedroni argues that it is not the most general approach.

The derivation of sequential limits is important because it permits control over misbehaving variables that are associated autocorrelation as T approaches infinity, it follows that the

construction of the limit $N \rightarrow \infty$ for cross-sectional data is simplified (Breitung, and Pesaran 2006;2008). Therefore we can associate properties of heterogeneity with the standard sum of stochastic variables $\sum_{i=1}^N R_i$ with reference to one serial correlated variable considered to contain conditional long-run variance of the differenced data L_{11i}^2 . Also as $T \rightarrow \infty$ the sequential limit theorem allows one to focus only on the first-order terms of the limit for the T-dimension, because before averaging higher orders are discarded over the cross-sectional dimension-N. These properties allow computation of the limit for the panel, although this property may pose problems in some cases, as it may not be able to show the need for controlling the relative expansion rate for the two dimensions, especially in the situation of a more general limit. The relative expansion rate also functions as an important indicator for small sample properties of the statistic for a panel consisting of different dimensions of N and T (Kao, 1999). Let us consider a group of statistics derived from pooling residuals from a regressed regression within the demission of the panel, also a regression obtained from pooling residuals from the between estimate of the panel. The idea is to hypothesize cointegration on both cases individually for each unit in the panel, then it follows that we pool together the resulting residuals in computing the test for no cointegration.

We can estimate $y_{it} = \alpha_i + \delta_i t + \beta_i X_{it} + e_{it}$ as a suggested cointegrating relationship for each unit in the panel separately, it follows that an intercept or trend is included if it warrants the model, to get the corresponding residuals \hat{e}_{it} . The manner with which pooled residuals are estimated is heterogeneous among many statistics(Kao, 1999). Pedroni (1999), defines pooled residuals as the panel and Group Mean Cointegration Statistics for Heterogeneous Panels. In the description of this statistic we allow $\tilde{e}_{it} = (\hat{e}_{it}, \hat{e}_{it-1})'$, $A_i = \sum_{t=1}^T \hat{e}_{it} \hat{e}_{it}'$ where \hat{e}_{it} the estimated residuals based regression, model (1). The statistic for testing the null hypothesis of no cointegration is described as

$$Z_{\hat{v}_{NT}} \equiv \left(\sum_{i=1}^N A_{22i} \right)^{-1}, \quad Z_{\hat{\rho}_{NT-1}} \equiv \left(\sum_{i=1}^N A_{22i} \right)^{-1} \sum_{i=1}^N (A_{21i} - T \hat{\lambda}_i)$$

$$Z_{\hat{t}_{NT}} \equiv \left(\sum_{i=1}^N A_{22i} \right)^{-\frac{1}{2}} \sum_{i=1}^N (A_{21i} - T \hat{\lambda}_i),$$

$$\tilde{Z}_{\hat{\rho}_{NT-1}} \equiv \sum_{i=1}^N A_{22i}^{-1} (A_{21i} - T \hat{\lambda}_i), \quad \tilde{Z}_{\hat{\rho}_{NT-1}} \equiv \sum_{i=1}^N (\hat{\sigma}_i^2 A_{22i})^{\frac{1}{2}} (A_{21i} - T \hat{\lambda}_i),$$

Let $\hat{\mu}_{it} = \hat{e}_{it} - \hat{\rho}\hat{e}_{it-1}$, $\hat{\lambda}_i = T^{-1} \sum_{s=1}^K w_{sK} \sum_{t=s+1}^T \hat{\mu}_{it} \hat{\mu}_{i,t-s}$ alternatives for lag window $w_{sK} = 1 - \frac{s}{(1-K)}$, $\hat{s}_i^2 = T^{-1} \sum_{t=2}^T \hat{\mu}_{it}^2$, $\hat{\sigma}_i^2 = s_i^2 + 2\hat{\lambda}_i$, $\hat{\sigma}_{NT}^2 \equiv N^{-1} \sum_{i=1}^N \hat{\sigma}_i^2$, and $\hat{L}_{11}^2 = N^{-1} \sum_{i=1}^N \hat{L}_{11i}^2$ if $\hat{L}_{11}^2 = \hat{\Omega}_{11i} - \hat{\Omega}'_{21i} \hat{\Omega}_{22i}^{-1} \hat{\Omega}_{21i}$ so that $\hat{\Omega}_i$ is consistent estimator of Ω_i (Breitung, and Pesaran 2006).

Computation of the first three statistics is based on pooling data within dimensions of the panels, this implies the test statistics are computed through summing up individually the numerator and denominator terms producing an unusual time series statistic (Pedroni, 2001; 2004). In practice a mean variance ratio statistic, different to the explained pooled panel cointegration test statistic can be computed, and it is found to be concentrated by the other in two components of the small size properties. The arguments made above are explaining nonparametric treatment of troubled parameters, although we should take note that nuisance parameters can also be thought of as being parametrical for panel and group mean statistics. The limit distribution proposed earlier is still applicable and it will be used in the parametric treatment of the form of panel and group mean augmented Dicky-Fuller (ADF) statistics. Pedroni (2004) made a proposition and applied the limit distribution for those test statistics for the null of no cointegration, let the indexes Θ , $\tilde{\Theta}$ and Ψ , $\tilde{\Psi}$ respectively denote a finite mean and the covariance of the corresponding vector Brownian motion function. The proposition made shows that when adequate values of T and N standardize the test statistic, it follows that an asymptotic distribution hinges only on parameters that are known generated by Θ , $\tilde{\Theta}$ and Ψ , $\tilde{\Psi}$ (Kao, 1999).

To test for the null hypothesis of no cointegration for heterogeneous panels from an asymptotic distribution of the residual based test. Let indexes Θ and Ψ denote vectors of mean and covariance of a function of a Brownian motion $\mathbb{T} \equiv (\int \mathfrak{h}^2, \int \mathfrak{h} d\mathfrak{h}, \tilde{\zeta} \tilde{\zeta}')$, when $\tilde{\zeta} \equiv V W' (\int W W')^{-1}$, where $\mathfrak{h} \equiv V - \tilde{\zeta} W$, while $\Psi_{(l)}$, $l = 1, 2, 3$ relating to $l \times l$ upper submatrix. If we let $\tilde{\Theta}$ and $\tilde{\Psi}$ describe Brownian motion $\mathbb{T}' = (\int \mathfrak{h} d\mathfrak{h} (\int \mathfrak{h}^2)^{-1}, \int \mathfrak{h} d\mathfrak{h} ((1 - \tilde{\zeta} \tilde{\zeta}') \int \mathfrak{h}^2)^{-1/2})$. Then we test the null hypothesis of no cointegration of the asymptotic distribution of residual based test, the statistic is given by

$$T^2 N^{3/2} Z_{\hat{\nu}_{NT}} - \Theta_1^{-1} \sqrt{N} \Rightarrow N(0, \Phi'_{(1)}, \mathfrak{q}_{(1)}, \Phi_{(1)}),$$

$$T \sqrt{N} Z_{\hat{\rho}_{NT-1}} - \Theta_2 \Theta_1^{-1} \sqrt{N} \Rightarrow N(0, \Phi'_{(2)}, \mathfrak{q}_{(2)}, \Phi_{(2)}),$$

$$TN^{-1/2}\hat{Z}_{\hat{\rho}_{NT-1}} - \hat{\theta}_1\sqrt{N} \Rightarrow N(0, \tilde{q}_{11}),$$

$$N^{-1/2}\hat{Z}_{\hat{\epsilon}_{NT}} - \hat{\theta}_1\sqrt{N} \Rightarrow N(0, \tilde{q}_{22}),$$

where T, N sequences get larger given $\Phi_{(l)}$ and is provided as

$$\Phi'_{(1)} = -\theta_1^{-2}, \Phi'_{(2)} = (-\theta_2\theta_1^{-2}, \theta_1^{-1}), \text{ it follows } \Phi'_{(3)} = \left(-\frac{1}{2}\theta_2\theta_1^{-2}, \theta_1^{-\frac{3}{2}}(1 + \theta_3)^{-\frac{1}{2}}, \theta_1^{-\frac{1}{2}}(1 + \theta_3)^{-\frac{1}{2}}, -\frac{1}{2}\theta_2\theta_1^{-2}, \theta_1^{-\frac{3}{2}}(1 + \theta_3)^{-\frac{1}{2}} \right).$$

When a panel cointegration test is done to determine the existence of a long run relationship in panel data, it is possible to find more than one cointegrating relationship in an equation (Aguirre and Calderon, 2005). In this section of the study, cointegration was tested using Johansen's cointegration for panel data, and the study uses trace statistics to determine cointegration across heterogeneous panels. The statistic assumes that there is cross-sectional independence. The trace statistic is standardized with the average of N individual trace statistic in addition. The trace statistic is asymptotically distributed and it follows a normal distribution with mean 0 and variance 1. To test for the existence of cointegration on residual, a cointegrating equation is estimated and we allow for the differences in the cointegrating vectors and short run variations between countries (Breitung, and Pesaran 2008).

$$RER_{it} = f(LNOPEN, LNGEX, LNGDPPCAP, BUDG, LNRESY)$$

This means individual countries are characterised by the relationship with reference to RER_{it} , and $LNOPEN_{it}$, $LNGEXP_{it}$, $LNGDPPCAP_{it}$, $BUDG_{it}$ and $LNRESY_{it}$. Where the index e_{it} denotes white noise innovation, the slope of the cointegrating coefficient is allowed to vary from country to country and to be none one. This is because Pedroni (2004), Pedroni (1999) and Aguirre and Calderon (2005) all assume that there is cross sectional independence across countries. If we have cointegration between these countries it means these variables have a point of equilibrium in the long run, and it follows that we can go on to estimate exchange rate misalignment and determine if currencies are undervalued or overvalued.

We run a residual based regression to compute an ADF based on pooled mean group. The critical values used for this cointegrating test are those computed by Pedroni (2004) and it is shown above that they allow one to run a test based on residuals instead of testing the relationship using a true relationship. The null hypothesis is that there is no cointegration.

Large and negative values mean that residuals are stationary and therefore it follows that we reject the null hypothesis of no cointegration, and choose the alternative hypothesis of non-stationary residuals meaning there is cointegration on the variable.

Table 4 below presents the results for Pedroni cointegration test using the Pedroni's (2004) cointegration test conducted on the cointegrating regression shown below using weighted estimates; four tests out of seven are not rejecting the null hypothesis of no cointegration. The ADF which uses a residual based regression test to determine the relationship explained above is also not reducing significantly at all conventional levels of significance. We can see that there is a cointegration relationship across the panels and so this leads to estimating the cointegrating relationship:

$$RER=f(\text{OPEN}, \text{LNGEXP}, \text{LNGDPPCAP}, \text{BUDG}, \text{RESY})$$

Table 4 Pedroni Conitegration test

Pedroni cointegration test				
	Statistic	Statistics with trend	weighted within-dimension	between-dimension
Panel v-Statistic	-1.98	-2.95	Yes	
Panel rho-Statistic	1.79	2.81	Yes	
Panel PP-Statistic	-3.7***	-3.23***	Yes	
Panel ADF-Statistic	-3.07***	-3.40***	Yes	
Group rho-Statistic	1.24	3.47		Yes
Group PP-Statistic	-2.05***	-4.37***		Yes
Group ADF-Statistic	-0.54**	-2.33***		Yes

Pedroni cointegration test. Null hypothesis: There is no cointegration. *, **, *** denote significance at 10%, 5% and 1%.

3.5 A panel bootstrap cointegration test

Common cointegration tests, at times can produce inadequate results unless the length of the time series is large. In this section I will test cointegration by using a method advocated by Westerlund and Edgerton (2007) which is based on the widely popular Lagrange multiplier test, which contains a null hypothesis of no cointegration. The test under asymptotic theory commonly produces an inadequate approximation to empirical test distribution. Employing bootstrapping techniques improves the test's performance. We implement the bootstrap to achieve structural independence of data, over time and across units, if properly replicated. This will be done in this thesis by adopting the approach that approximates dependence of the

time series in equilibrium errors, employing a finite order of autoregressive model (Coulibaly and Gnimassoun 2013). The preservation of the cross-sectional dependence on bootstraps drawn are constructed from the combined empirical distribution of regress errors.

Assumptions underlying the model

Suppose a scalar variate y_{it} which is a function of

$$y_{it} = \alpha_i + x'_{it}\beta_i + z_{it} \quad (30)$$

Where $t = 1, \dots, T$ and $i = 1, \dots, N$ the indexes denote the time series and cross-sectional units, respectively. The vector x_{it} contain N dimensions and has regressors that are assumed to follow a pure random walk process. The error term z_{it} is assumed to be given by

$$z_{it} = u_{it} + v_{it} \quad \text{where } v_{it} = \sum_{j=1}^t \eta_{ij} \quad (31)$$

Where η_{it} is i.i.d process with a mean equalling to zero and the variance $\text{var}(\eta_{ij}) = \sigma_i^2$, and the vector $w_{it} = (u_{it}, \Delta x'_{it})'$ is a linear process satisfying;

$$w_{it} = \sum_{j=0}^{\infty} \alpha_{ij} e_{it-j}, \quad (32)$$

Where e_{it} is i.i.d errors with a mean of zero across t , while the indexes α_{ij} are assumed to fulfil common summable requirements. Because α_{ij} differs across i , this model gives way for a completely heterogeneous autocorrelation structure (Westerlund and Edgerton 2007). To allow for cross-sectional dependence, we permit the staked time series vector $e_t = (e'_{1t}, \dots, e'_{Nt})'$ to allow for a positive precise covariance matrix $\text{var}(e_t) = \Omega$ (see Coulibaly and Gnimassoun, 2013). This is to test for the null hypothesis of cointegration against the alternative of no cointegration, which can be expressed as $H_0: \sigma_i^2 = 0$ for al i against $H_1: \sigma_i^2 > 0$ for a certain i . When we have cross-sectional independence, the hypothesis can be tested by employing the following test statistic

$$LM_N^+ = \frac{1}{NT^2} \sum_{i=1}^N \sum_{t=1}^T \hat{\omega}_i^{-2} S_{it}^2$$

Where S_{it} denote the partial sum process of \hat{Z}_{it} , the completely transformed estimate of Z_{it} and $\hat{\omega}_i^2$ is the estimated long-run variance of u_{it} which depends on Δx_{it} . (discuss conditions from McCoskey and Koa, (1998)). This statistic is a right-tailed Lagrange multiplier hypothesis test of H_0 against H_1 . Westerlund and Edgerton (2007) describe this test as the most powerful unchanging and unbiased test for unit moving average root in Z_{it} . The sing \Rightarrow indicate convergence in distribution.

$$\sqrt{N}(LM_N^+ - E(LM_N^+)) \Rightarrow N(0, var(LM_N^+)), \text{ as } T \rightarrow \infty, N \rightarrow \infty$$

The concern with the results is that they depend on Ω which is block-diagonal, making e_{it} which is cross sectional independent, this seldom hold in applications. Westerlund and Edgerton (2007) argue that other studies found the test to be very sensitive to autocorrelation even in situations where error are cross-sectionally independent. Basing the inferences on the theory of critically normal critical values can therefore be largely misleading in small sample sizes. The proposed method by Coulibaly and Gnimmassoun (2013) of dealing with this problem is through bootstrap test. They suggest the sieve bootstrap scheme which is encouraged by the fact that, when w_{it} meet all the conditions we have discussed earlier, it follows autoregressive representation.

$$\sum_{j=0}^{\infty} \phi_{ij} w_{it-j} = e_{it} \quad (33)$$

The reason of approximating $\Delta y_t = \phi y_{t-1} + \varepsilon_t$ using the model of order p_i , a finite constant, which results to the formation of residual-based resampling plan. When we allow p_i to rise at a particular rate with T , the process in $\Delta y_t = \phi y_{t-1} + \varepsilon_t$ will be exactly equal asymptotical (check Chang and Park (2003)). To bootstrap initially we need to estimate ϕ_{ij} in equation (33) by using $\hat{\omega}_{it} = (\hat{z}_{it}, \Delta x'_{it})'$ rather than w_{it} and p_i lags. Therefore residuals can be computed as follows

$$\hat{e}_{it} = \sum_{j=0}^{P_i} \hat{\phi}_{ij} \hat{w}_{it-j} \quad (34)$$

Given \hat{e}_{it} , we can form a vector $\hat{e}_{it} = (\hat{e}'_{1t}, \dots, \hat{e}'_{1N})'$. To make certain that the autoregression in equation (34) is always invertible, we use the empirical Yule-Walker to help choose parameter estimate $\hat{\phi}_{ij}$ (look at Lutkepohl (2005)). At this stage this becomes very important

in a situation where equation (34) is not invertible, the generated bootstrap sample in the next step is to obtain the random sample of e_t^* derived from the empirical that weighs $1/T$ on individual residuals that are centred $\hat{e}_t - \frac{1}{T} \sum_{j=1}^T \hat{e}_j$. Then w_{it}^* is generated from e_t^* and w_{it}^* from e_t^* in a recursive manner through employing equation (34) but replacing w_{it}^* and e_t^* replacing \hat{w}_{it} and \hat{e}_{it} in this respective order

Then the next step is to segment w_{it}^* into $w_{it}^* = (z_{it}^*, \Delta x_{it}^*)'$ and obtain the bootstrap samples x_{it}^* and y_{it}^* through application of this recursion;

$$y_{it}^* = \hat{\alpha}_i + \hat{\beta}_i + z_{it}^*, \text{ where } x_{it}^* = \sum_{j=1}^t \Delta x_{ij}^* \quad (35)$$

where indexes $\hat{\alpha}_i$ and $\hat{\beta}_i$ are completely changed estimates of α_i and β_i (Chang and Park 2003). Upon getting bootstrap sample y_{it}^* and x_{it}^* , the bootstrap test statistic obtained is analogous to the previous sample. When this process is repeated S times we get a bootstrap distribution that gives us the test statistic. We then conduct a one-sided nominal level test (if we test for 5% critical values are calculated as the lower 5th percentile of the bootstrap distribution, this approach is an alternative to conventional critical values).

To be able to analyse small sample properties of the bootstrap test by using Monte Carlo simulation analysis, employing the data generating process in equation (30) defining error term by the following function

$$z_{it} = \lambda_i F_t + e_{it} \quad \text{with } e_{it} = \rho e_{it-1} + u_{it}, \quad (36)$$

Where u_{it} exhibits the Brownian motion $u_{it} \sim N(0,1)$, $\Delta x_{it} \sim N(0,1)$ and $F_t \sim N(0,1)$ which is a common factor that brings about cross-sectional dependence. Through increasing the factor λ_i the extent to which there is dependence is obtained, since increasing factor λ_i we give way to the presence of heterogeneity between units, these cross sectional units are drawn once for every individual replication from a normal distribution. The mean and variance of λ_i is being used to control $\phi = \text{cor}(z_{it}, z_{jt})$ and $\sigma = \frac{\text{var}(\lambda_i F_t)}{\text{var}(z_{it})^2}$. To make matters easy, suppose that there is a single repressor that has α_i and β_i drawn from $U(1, 2)$. In this setup we test the null hypothesis $H_0: \rho < 1$ against the null hypothesis of $H_1: \rho = 1$ (Westerlund and Edgerton 2007;).

Westerlund and Edgerton (2007) attests that when data is generated for 1000 replications for N cross-sectional observations, and with $T + 50$ time-series observations, for the first 50 observations for each, cross-sectional units are removed to allow for the reduction of the effect of the initial conditions, which were all initially set to zero. The validity of the bootstrap test must be consistent with the theoretical requirements which requires that the autoregressive order should be allowed increase along with T , we define p_t as $4(T/100)^{2/9}$. This theoretical rule is used to implement a completely modified estimator of equation (20), which needs a choice of a Kernel bandwidth. The interest is in comparing small sample characteristics of the bootstrap test against those of the asymptotic test proposed by McCoskey and Kao (1998), we analyse differences between different values of ρ, ϕ and σ . We also want to determine if these characteristics vary for different sample sizes as they increase.

Coulibaly and Gnimassoun (2013) are convinced that the bootstrap test is characterised by adequate size accuracy, they arrived at this determination after running a series of tests and they found this to be true for all experiments they conducted. The asymptotic test is usually overestimated, more especially in situations where the time series component has autocorrelation. It is argued that a distortion for the asymptotic test has a tendency of accumulating as N grows.

Table 5 presents the results for the Westerlund cointegration test that accounts for the dependence between cross-sectional units are it also accounts for structural breaks using Westurlund and Edgerton (2007) test. Upon testing for cointegration, three out of four tests indicate the existence of cointegration between variables. The Gt test indicates cointegration at 5%, while both Ga and Pa indicate cointegration at 10%.

Table 5 Westerlund Cointegration test, Westerlund (2007)

Westerlund cointegration test	
statistic	Z-value
Gt	-2.547**
Ga	1.867*
Pt	-0.969
Pa	0.677*

Westerlund cointegration test. Null hypothesis is no cointegration. .*, **, *** denote significance at 10%, 5% and 1%.

Chapter 4

4.0 Empirical Analysis and Discussion

On this section of the paper I estimate exchange rate misalignment then discuss finding with reference to economic theory. This study continuous to estimate values of real exchange rate misalignment and the index of real exchange rate misalignment; this is extensively practiced in economic literature. We make the assumption that imposes real exchange tare misalignment at any period of time is described $\log e_t = \hat{\alpha} + \hat{\beta}' F_t$, (Kamar and Naceur, 2007)where index F denote the long run fundamentals and of estimation of coefficients of relevant parameters, then Hodrick- Prescott filter procedure is applied to decompose time series fundamentals into parameters (\tilde{F}) and transitory ($F - \tilde{F}$) components. The next step is to compute real exchange rate equilibrium; $\log \tilde{e}_t = \bar{\alpha} + \hat{\beta}' \tilde{F}_t$, index $\hat{\beta}'$ represents estimated coefficients and $\bar{\alpha}$ denotes the intercepts which correspond to each country. It only significant real exchange rate misalignment defined by $rem_t = (\log e_t - \log \tilde{e}_t) \times 100$ with negative value of real exchange rate indicate overvaluation and positive values denote that real exchange rate is undervalued. But we have to acknowledge that computation of real exchange rate is subject to criticism.

The graph measures the logarithm of exchange rate misalignment on the vertical axis as well as the log of per capita GDP growth rate. On the horizontal axis the curve measures time/years. For this thesis the main focus is on the mean group estimator, since real exchange rate dependence is found within countries. Negative values of real exchange rate misalignment relates to currency overvaluation, and positive values of real exchange rate misalignment means the currency is undervalued. The reason is that real exchange rate misalignment is calculated by actual real exchange rate, subtract equilibrium real exchange rate. Annual data is used for all time-series and the data used in this study is obtained from the World Bank development indicators.

Botswana is showing a downward trend for real exchange rate misalignment indicating a steady and constant move away from equilibrium exchange rate. Real exchange rate misalignment is negative, implying that from 1995-2012 the currency in Botswana was overvalued. Economic growth does not seem to be influenced or responding to variations occurring around real exchange rate disequilibrium. Economic growth rate is volatile in early years as real exchange rate deviates further away from equilibrium exchange rate. In 1998,

2001 and 2009 exchange rate misalignment and per capita GDP growth rate are (-3.47;-1.33), (-4.82;-1.33) and (-4.62;-9.09) respectively. In all the years preceding the year when real exchange rate deviated from equilibrium exchange rate, economic growth was lower in the following year. This implies that there is a positive weak correlation between real exchange rate misalignment and economic growth; as real exchange rate misalignment gets larger, economic growth fluctuates in the region of positive values. The results analysed reveal weak positive correlation which does not support Rodrik's (2008) claims that undervaluation of a developing country's currency increases economic growth.

In Botswana after 2001, economic growth becomes positive for the following 7 years, and then increases to a high level of 7.31 percentage points. Real exchange rate misalignment increases even more on the other hand as economic growth increases by higher rates and falls again in 2009. The decline in economic growth rate seen in 2009 can be associated with the aftermath of the global financial crisis. Soon after 2009, real exchange rate continues to move away from equilibrium exchange rate, and per capita GDP growth rate shoots up to its highest level in 17 years, to 7.34 percentage points. It appears that real exchange rate misalignment does not have a clear and direct effect on economic growth for Botswana.

Lesotho's real exchange rate misalignment is not as large as the one for Botswana. Lesotho's real exchange misalignment is negative, before moving away from equilibrium real exchange rate until 2002, which is a trough of -2.05 percentage points. From that point onwards the real exchange rate began to move backwards towards equilibrium real exchange rate. Initially, as the exchange rate's overvaluation increased, economic growth rate decreased until it reaches negative growth rate in 1999 of -0.58 percentage points. In the following year, GDP growth increased to a positive growth rate, before falling again in 2002 to a negative growth value of -0.21 percentage points. In the years following this year, real exchange rate misalignment fell towards equilibrium real exchange rate, while economic growth increased and fluctuated with high positive values. From 2004 to later years, real exchange rate misalignment and GDP growth exhibited similar movements, as real exchange rate misalignment appears to have an effect on economic growth rate in the case of Lesotho. It appears that if real exchange rate misalignment fluctuates around equilibrium, real exchange rate economic growth increases and follows a similar pattern to that displayed by real exchange rate. This is consistent with the findings made by Aghion et al (2009).

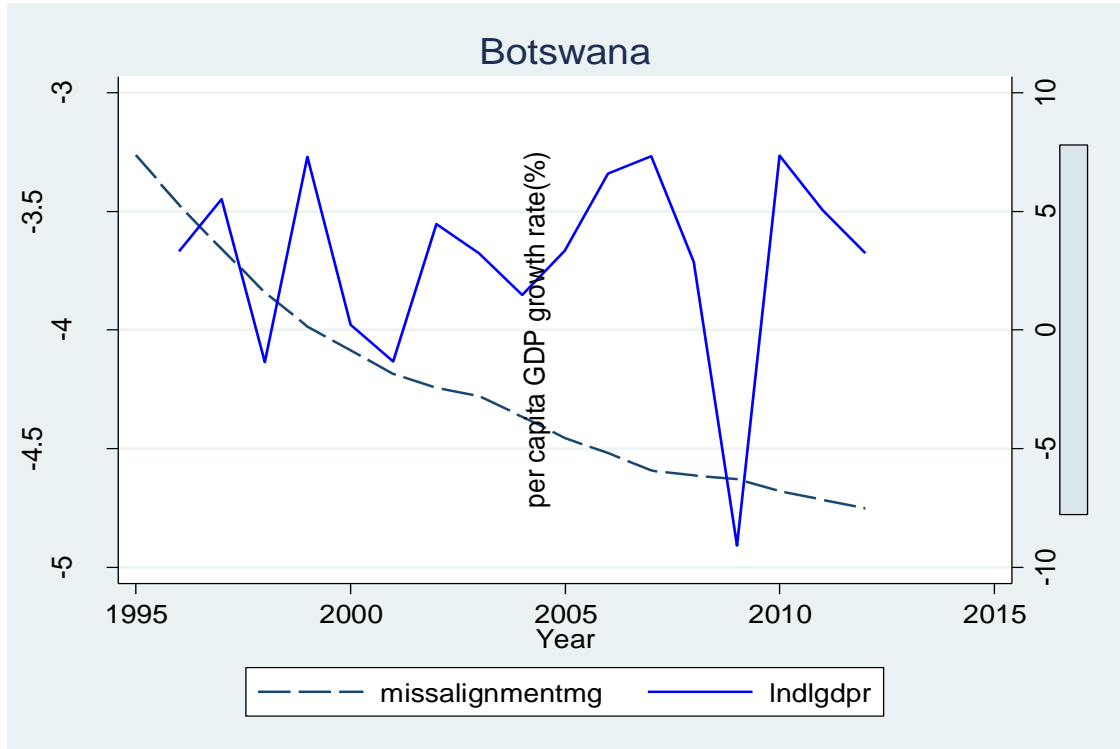
For Namibia, real exchange rate misalignment behaves differently than for the two countries discussed above. Real exchange rate misalignment is negative, but at a level closer to equilibrium exchange rate when compared to the other two countries, and it moves further away from equilibrium exchange rate (i.e. real exchange rate becomes more overvalued). Throughout the period, real exchange rate misalignment and economic growth have the same movement; when exchange rate misalignment becomes more negative, economic growth rate declines. But when real exchange rate misalignment declines (i.e. moves back to equilibrium exchange rate) economic growth increases. However, after 2002 when the log of real exchange rate misalignment was -3.17 percentage points, the highest overvaluation of the currency in 17 years, economic growth rate in Namibia increased to its highest in 17 years, which is 10.47 percentage points. Following that year, real exchange rate misalignment declined and the currency depreciated, in 2004 when real exchange rate misalignment is close to equilibrium level that in the previous year in 2005 economic growth declines by 1.32 percentage points. The logarithm of real exchange rate misalignment is fairly constant from 2004 to 2006, after that it increases. Therefore the real exchange rate becomes more overvalued and economic growth rate follows the same pattern as shown in the graph for Namibia. In 2008 real exchange rate misalignment deviates from equilibrium by -3.08 percentage points; in the following year (2009), GDP growth declines to a negative economic growth of -2.63 percentage points. In the following year real exchange rate misalignment moves back towards equilibrium exchange rate, and economic growth behaves in the same way. In the early years of the analysis, economic growth appears to be following the same pattern as exchange rate misalignment, such that when real exchange rate misalignment moves towards equilibrium, exchange rate and economic growth increases. However, when real exchange rate misalignment deviates from equilibrium, economic growth declines. In the intermediate years the response of economic growth rate is different; when real exchange rate misalignment increases, economic growth rate also increases.

South Africa's real exchange rate misalignment increased for the whole period of 1995 to 1996, then moved back towards equilibrium real exchange rate from 1997 to 1998. After these years real exchange rates became overvalued and deviated from equilibrium real exchange rate. In this time period economic growth increased by 1.99 percentage points, and then declined as real exchange rate misalignment increased by -1.84 percentage points. In the following years, after a slight depreciation of the real exchange rate, misalignment per economic growth improved and started fluctuating in the region of positive values. In 2002

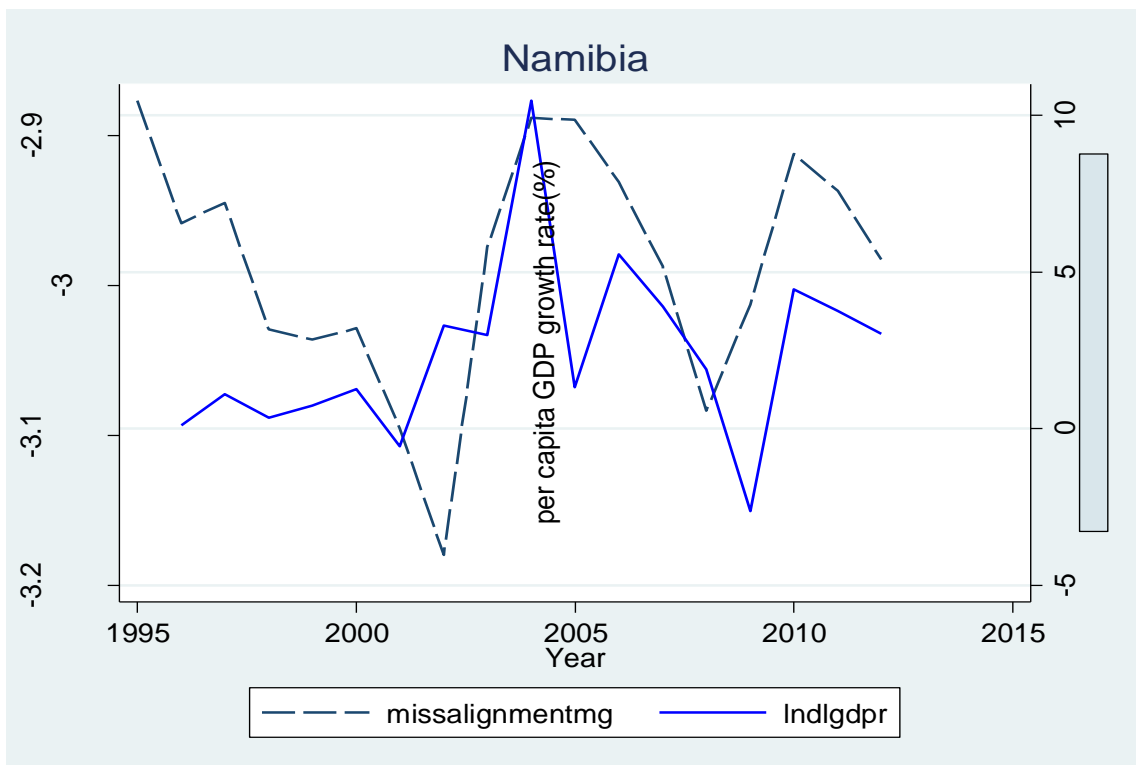
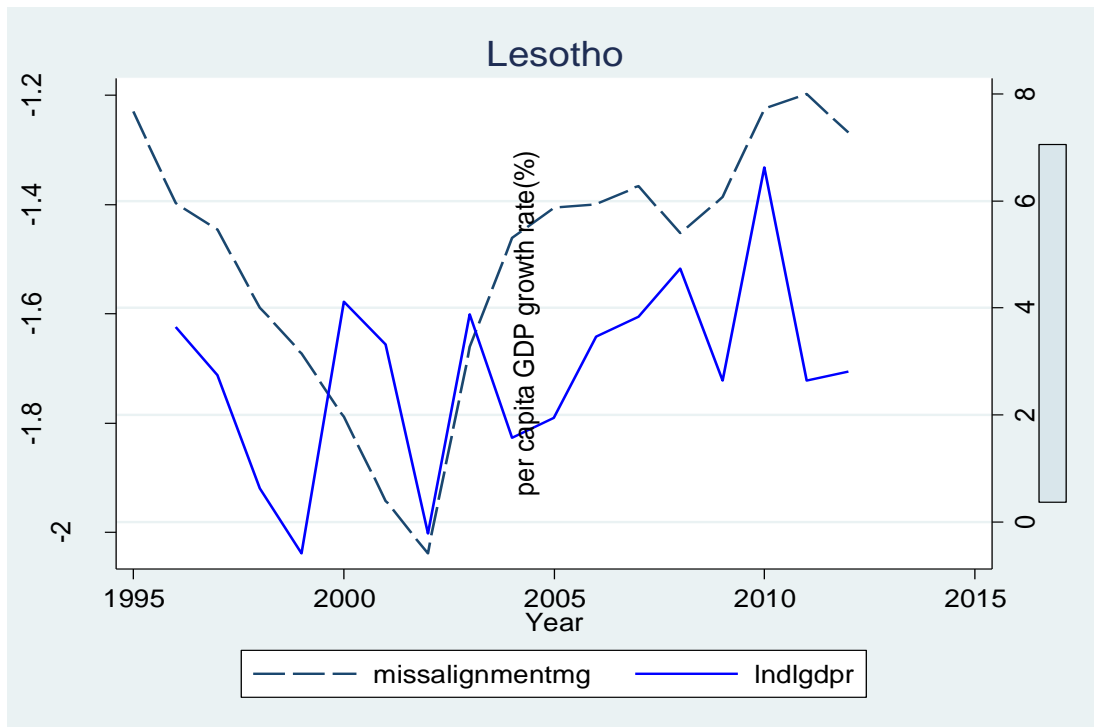
real exchange rate misalignment declined to -3.42 percentage points, and in the following year real exchange rate misalignment increased, economic growth increased and became stable at a rate above 4 percentage points for the time period of 2005 to 2007. It is worth noting that in this time interval, the real exchange rate became more overvalued, and as real exchange misalignment increased, further economic growth declined. In 2009, real exchange rate misalignment fell by a slight margin while economic growth grew by -2.61 percentage points. Following this slight decrease in real exchange rate misalignment, in 2010 economic growth rose to a positive growth of 1.69 percentage points. Thereafter, real exchange rate misalignment increases, while at the same time economic growth increases between 2010 and 2011, and then falls. Real exchange rate misalignment variation affects economic growth with a lag of 12 months or slightly more.

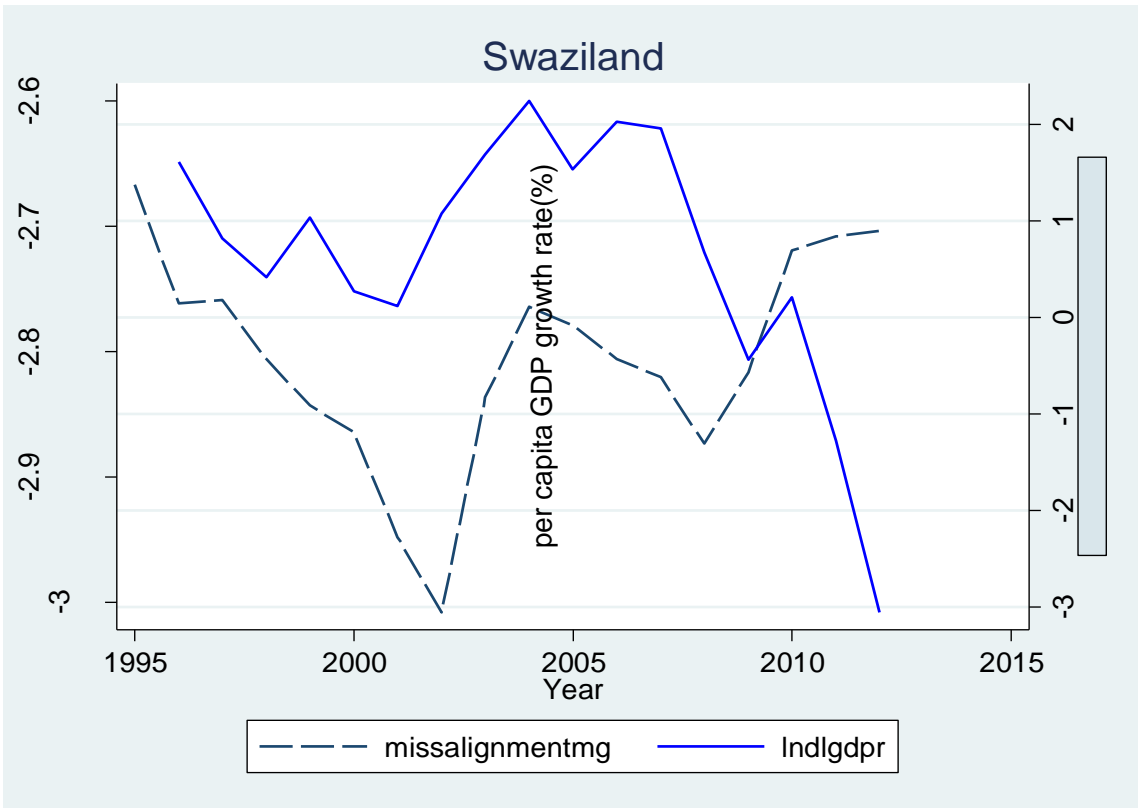
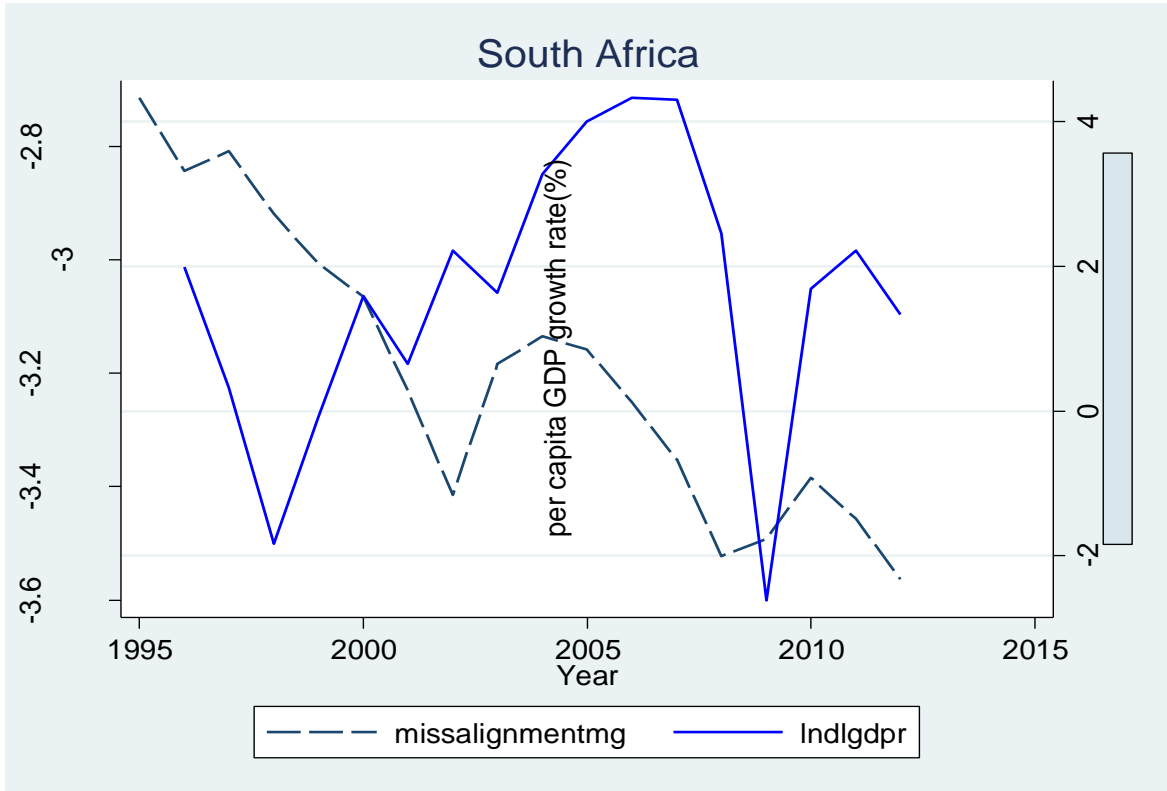
Real exchange rate misalignment for Swaziland keeps increasing and stabilizes in 1999 to 2000, and then it increases until 2003. Thereafter real exchange rate misalignment declines moves towards real exchange rate equilibrium. From 2004 to 2008 real exchange rate misalignment moves away from equilibrium exchange rate, and continues from 2008 until the end of the period of analysis. For the duration of 1996 to 1998, economic growth rate falls as real exchange rate becomes more overvalued and then the economy expands from 1998 to 1999. The increase of real exchange rate misalignment does not seem to induce an increase in economic growth. In 2005 real exchange rate misalignment reaches its trough and economic growth is also at a turning point. As real exchange rate misalignment falls, economic growth rate increases but with economic growth delays. In 2007, real exchange rate misalignment increases and economic growth declines, although in 2008 misalignment reaches a trough but economic growth continuous to fall until 2009. Economic growth is responding with a lag and that is influenced by variations occurring around real exchange rate equilibrium. In 2010 real exchange depreciates at a fairly steady rate and economic growth falls in the following years. Swaziland is the only country that has economic growth, experiencing a double deep in recession after the global financial meltdown. SACU countries seem to have responded late to the global financial crisis, but what is important is that they were able to adjust back to positive economic growth rate.

Misalignment_MG and per capita Growth rate



missalignmentmg





4.1 Estimation Results

Following Magyari (2008), Rodrik (2007) and Razinand Collins (1997), we estimate the following regression:

$$grwth_rte_t = \gamma_0 + \gamma_1 grwth_rte_{t-1} + \gamma_2 pcgdpgap_{t-1} + \alpha_e erm + \beta' X_t + \xi_t$$

where $grwth_rte_t$ is a dependent variable denoting a differenced log of per capita GDP, capturing growth rate of the gross domestic product, it is an economic variable that is appropriate for use to capture economic growth rate. $grwth_rte_{t-1}$ denotes lag of economic growth rate of real GDP, $pcgdpgap_{t-1}$ denotes the first order lag of GDP output gap, which captures cyclical revision and the effect it has on economic growth. The GDP output gap is calculated as a percentage change of per capita GDP (which is used to model real gross domestic product) from its potential level of output, we cannot observe potential level of gross domestic product. Rodrik (2007) estimated it because he considered it to have policy implications. Potential levels of gross domestic product is estimated by employing Hodrick- Prescott filter, HP filter parameter λ , is set to 1600, as Magyari (2008) did following Hodrick- Prescott's (1997) suggestion.

It must be noted that the HP filter was developed to explain the post-war business cycle, and the HP parameter was set to 1600, so the concern is that this parameter might not be the same for other economies. Magyari (2008) also raises this caution. $erm_{i,t}$ (in some cases it is presented as $misalignment_MG$, if estimated by the MG estimator, or $misalignment_PMG$ if estimated by the PMG estimator) denote real exchange rate misalignment, it is estimated by following the procedure explained in the previous section. An additional vector of explanatory variables is represented by X_t in which it contains growth rate of the log of degree of openness which measures domestic structural policies and the growth rate of the terms of trade which measures the external environment. The model used to measure the economic growth rate is the GMM model. This model is able to control for the problem of endogeneity that might arise in explanatory variables used to explain the rate of economic growth. Since lagged variables are used, there is a strong possibility that the model might be suffering from endogeneity. OLS estimators would not be appropriate to use as it would generate inconsistent estimators.

. The variables used on this model are economic growth rate ($grwth_rte$), lagged growth rate as a proxy for initial growth ($grwth_rte(-1)$), per capita income ($pcgdpgap$), real exchange

rate misalignment (*rem*), growth rate of openness (*gwthop*) and growth rate of term of trade (*gwthtot*).

4.2 Estimating correlation between real exchange rate misalignment and economic growth in SACU

Model 1 real exchange misalignment (*rem*) is statistically insignificant at all conventional levels.. The initial growth is positive and significant at 1% level of significance. So countries with high initial growth rate tend to achieve higher growth rates in future. The growth rate of openness is significant at 1% and has a positive coefficient. So the larger the international trade, the higher the economic growth. The growth rate of terms of trade has a negative coefficient and insignificant at all conventional levels.

The model estimated above is the most robust model. A model that was not robust was estimated and it produced similar results to those produced by a robust GMM estimation. Indexes μ_i and λ_t denotes unobserved country-by-specific effects and time specific effects as fixed effects. Sargan's test of over-identification of restriction is computed to determine if specified instrumental restrictions are not collinear with error terms, as suggested by Arellano and Bond (1991). In computing this test, a J-statistic is estimated and a generated p-value is 0.35. Rodrik (2007), Arguirre and Caderon (2007), Razin and Collins (2005) and Magyari (2008) all dispute negative effects of overvaluation of the currency on growth, and positive effects of undervaluation on economic growth.

Models 2 and 3 provide the same results and are robust to serial correlation. In these models, only the initial growth rate and the per capita income are significant and positive; with the same interpretation that model 1. Model 4, while controlling for openness and terms of trade, has only the income per capita as the significant variable with the correct sign. Models 5, 6 and 7 are, unfortunately, not robust to serial correlation as shown by the test of over-identification and excluded for interpretation.

Thus, looking at all the models we can conclude that the exchange rate misalignment, computed using the MG estimator, is insignificant in explaining growth rates of these countries, even when we control for other variables.

GMM model specification

Variables	Model 1	Model2	Model 3	Model 4	Model 5	Model 6	Model 7
grwth_rte(-1)	0.36***	0.36*	0.36**	0.39	0.37*	0.36	0.39
pcgdpgap	0.78***	0.78***	0.78***	0.80**	0.78***	0.78***	0.81**
rem	0.18	0.18	0.13	-0.01	0.18	0.13	0.56
gwthopen	0.08**	0.08	0.07		0.08	0.07	
gwthtot	-0.03	-0.09			-0.03		
cons	1.77	1.78	1.61	1.22	1.77	1.01	0.43
Sargan test	78.16	78.16	77.84	73.1	78.16	77.89	75.36
	lag(2 2)	lag(2 2)	lag(2 2)	lag(2 2)	lag(2 4)	lag(2 4)	lag(2 4)
	robust	Robust	robust	robust			

*,(**)*** means significant at 10, 5 1 percent

5.0 Conclusion

The main objectives of this study are to establish the existence of macroeconomic policy coordination between the countries, compute the exchange rate misalignments in order to determine if currencies are overvalued or undervalued and analyse the impact of exchange rate misalignments on economic growth.

We use annual data from 1995 to 2012 and compute the REER using the BEER approach. After testing the unit root, the cointegration between real exchange rate and its fundamentals are investigated using the Westerlund test. The exchange rate misalignments are computed as the deviation of observed exchange rates from the REER using only significant long-run estimates interacted with the different detrended fundamentals. The last step examines the impact of exchange rate misalignments on economic growth. Due to the presence of endogeneity, we use the two system GMM approach.

The result show that Lesotho, Namibia, and Swaziland have macroeconomic coordination, their economic activity in all years under analysis exhibit similar behaviour because we found all countries to be cointegrated using Westerlund test, the test showed significant results for cointegration for all these countries. The other two countries South Africa and Botswana display macroeconomic coordination with each other for the entire sample period. Second paragraph to be related with your first objective SACU countries are dependent on each other, therefore, PMG estimator is used to estimate exchange rate misalignment, and the

exchange rate misalignment is overvalued for all these countries. Zehirum et al (2014) conducted a similar study using 12 SADC countries and they arrived to similar finding that exchange rate misalignment is overvalued, the graphs above show negative values. Empirical evidence reveals that exchange rate misalignment does not have impact on economic growth on all SACU countries, and there is no correlation between economic growth and exchange rate misalignment, the variable of exchange rate misalignment on economic growth is in significant meaning there relationship between economic growth and exchange rate misalignment. Currency overvaluation does not stimulate economic growth for SACU countries.

There should be policy hominization in SACU countries if these countries seek to establish economic and monetary union, South African is the most dominant economy among all countries in the SACU. However, it is one of the two countries have more overvalued currency, South Africa and Botswana must reduce their exchange rate misalignment to the level of Namibia, Lesotho and Swaziland, only then can these countries can launch a successful single currency union.

The contribution of this paper is that it established that there is dependence between SACU economies, which was found by testing for dependence using CIPS test. In later years, Lesotho, Namibia and Swaziland's competitiveness is moving in a similar direction. South Africa and Botswana are less competitive than other member states, but are closely related to each other as they need to coordinate their policies to prevent an unstable economic and monetary union.

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