The Real Exchange Rate of Oil Exporting Countries: An African Experience

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Abstract
Oil prices traditionally have been more volatile than many other commodity or asset prices since World War II and has have a lot implications on major macroeconomic variables such as inflation, exchange rate, money supply, capacity utilisation and economic growth among others. The goal of this paper is investigate the long run effects of real oil price and real interest rate differential on real exchange rate for a quarterly panel of 3 countries from 1980 to 20012. The modelling exercise follows three steps. In the first step, the paper investigates the integration properties of the data and finds them to be integrated of order one. In the second step, using several different panel cointegration tests, the paper finds evidence for cointegration among the three variables. In the third step, using pooled mean group estimator, the paper finds a positive and statistically significant impact of real oil price on real exchange rate for net oil importing countries, implying that increase in oil price leads to real exchange rate depreciation.
Keyword: Exchange rate; oil price; Cointegration

1.0 Introduction:

Global economy runs almost practically on oil. Sharp fluctuations in the oil price provoke
Substantial shifts in the wealth of nations especially the oil involved directly in trade of same.The large increase in oil prices since the beginning of the new millennium has been associated with the emergence of large current account imbalances across the globe. As the current oil shock has proved to be more persistent than expected, oil exporting countries have emerged as the group of countries with the largest current account surplus. This has prompted renewed interest in the economies of these countries and, in particular, their exchange rates as this determines how much gain they make from international trade,That the degree of nominal exchange rate flexibility of oil exporting countries is dependent on the behaviour of the real exchange rate over the long run, considering the possibility that this exchange rate may in turn be influenced by significant changes in the terms of trade. This without doubt, is a likely determinant of the real exchange rate.In oil exporting countries, the main driver of the terms of trade is the oil price.

High oil prices are again transforming oil-exporting economies. National Economies that were almost crashing when oil price hovered around $20s for most of the 1990s—and at risk of bankruptcy when oil dipped to $10 a barrel in 1998—are now booming. A new generation face of development is vivid in these countries, in Abuja, and in Luanda. Government coffers in oil-exporting economies are overflowing with the governments' (typically very large) cut from the oil windfall. Most oil-exporting economies now need an oil price of $40 a barrel to cover their import bill, including their bill for imported labour—up from $20 a barrel a few years ago. But with oil trading above $90 a barrel, they still have substantial sums available to invest in the rest of the world.

One feature of oil-exporting most oil exporting economies, though, has not changed to their tendency to peg to the dollar though some other peg to a basket of currency which comprise mostly of the dollar and the euro which is all beneficial as the two countries that that own these currencies are oil importing which indicate same pattern flow in there currency value as oil price per time.

The strong dollar contributed to the difficulties many emerging economies experienced when oil prices fell in the late 1990s. Falling government revenues led oil exporters to cut spending, draw down their external assets, run up their debts and in some case devalue and default. However, more recently, the weak dollar has made it more difficult for the oil exporters to adjust to the rise in the price of oil. After an initial period of surprising prudence, the oil-exporting economies are increasing government spending and investing heavily in large government-sponsored mega projects. The result is high inflation: Several smaller Gulf economies now have inflation rates well above 10 percent like in the case of Algeria between 1991and 1996 and a current condition in Gabon though quite fair in Nigeria.

The oil-exporting economies themselves have the most to gain from less rigid exchange rate as the world economy would.it is worthy of note that the large trade and current account surpluses of oil-exporting economies stem from high oil prices, not from any competitive edge from their undervalued currencies. But the large increase in their dollar holdings over the past years has helped to mask the consequences of the large US current account deficit. More important, global adjustment will remain more difficult than it needs to be so long as the currencies of many large surplus countries remain tightly tied to the currency of a large deficit country although pegging to the dollar allows for the importation of the United states' relatively stable monetary policy.

Work by the International Monetary Fund (IMF) indicates that a 100 percent increase in the real price of oil typically leads to a 50 percent real appreciation of the currencies of oil-exporting economies (Lee, Milesi-Ferretti, and Ricci 2007). This adjustment could come from a change in the exchange rate. Countries that allow their currencies to float—even with extensive management—would likely experience a nominal appreciation when oil is strong and a nominal depreciation when oil is weak (Frankel 2006). Countries that peg to the dollar or another currency could achieve a similar result through a one-off revaluation—or devaluation.
However, if the country's exchange rate remains fixed, the adjustment in the real exchange rate necessarily will come through changes in domestic prices. A rise in the price of oil implies a temporary rise in inflation; a fall in the price of oil implies a period of deflation. If an oil-exporting economy pegs to the dollar, the need for a change in domestic prices would be present even if the dollar holds steady relative to other currencies. But if the dollar falls relative to other currencies, pulling down the nominal exchange rate of the oil-exporting economies that peg to the dollar, the increase in inflation needed to generate the expected real appreciation goes up. Both the rise in the price of oil and the fall in the dollar put pressure on domestic prices.
It is worthy of note to say that government's oil export revenue will change the way the windfall is distributed. If the country's currency is pegged to the dollar, the government initially captures the entire windfall through a rise in its revenues. If, by contrast, the country's currency rose along with the price of oil, the government's local-currency revenue windfall would be smaller, but the external purchasing power of all local salaries would rise.
This study uses real oil price as a proxy of the terms of trade and questions the effect of oil price fluctuations on the real effective exchange rate of three oil exporting African countries: Nigeria, Angola and Algeria. Since these countries adopt different exchange rate regimes
Bearing in mind the level of reliance these countries have in their oil sector, oil exports accounts for around ninety percent of total exports in Gabon and contributes 45% to the nation's gross domestic product. In Nigeria,The oil and gas sector accounts for about 35 per cent of gross domestic product, and petroleum exports revenue accounts for about 70 per cent of total exports revenue For Algeria, the oil and gas sector contributes 35% to her gross domestic product and two-thirds of her total export. However, these three countries recorded a significant contribution to oil production both in African and across the globe.

2.0 Empirical Review:

The potential role of oil price variations in deriving terms of trade movements and impacting on exchange rates has already received much attention in literature. Several studies have been carried out without totally eliminating the imprecision of the effect of oil price variation on exchange rate for both oil importing and exporting countries alike. Some claim the existence of a relationship between oil prices and exchange rates in both developed oil importing and oil exporting countries. From a theoretical perspective, Krugman (1983) and Golub (1983) were the first to develop models in which shifts in oil prices generate wealth transfer effects and portfolio reallocations, leading to adjustments in exchange rates to clear asset markets. As regards the bilateral exchange rates between two or more oil importing countries, the relative propensity to import oil and their respective bilateral trade deficits against oil producing countries are the key variables in explaining whether a rise in the oil price will lead to an appreciation or depreciation of the currency.
Lastrapes (1992) by using the Blanchard and Quah (1989) decomposition, find that much of the variance of both real and nominal exchange rates from a number of countries over both short and long horizons is due to real shocks. The conclusions from the structural time-series literature therefore seem to be robust to both decomposition methods and currencies. This has led some to suggest that an unidentified real factor may be causing persistent shifts in real equilibrium exchange rates. also using the Blanchard identification strategy, Clarida and Gali (1994) estimated the share of exchange rate variability that is due to different shocks by using quarterly US–Canada, US–Germany, US–Japan, and US–UK real exchange rate data from 1974:Q3 to 1992:Q4. They find that real shocks can account for more than 50% of the variance of real exchange rate changes over all time horizons. Different sources of real shocks havebeen investigated in Zhou (1995). Amongmany sources of real disturbances, such as oil prices, fiscal policy, and productivity shocks, it has been shown that oil price fluctuations play a major role in explaining real exchange rate movements.
Moreover, Chaudhuri and Daniel (1998) investigate 16 OECD countries and find that the non-stationary behavior of US dollar real exchange rates is due to the non-stationary
behavior of real oil prices. Similar results are obtained by Amano and Norden (1998a, b). By using data on real effective exchange rates for Germany, Japan, and the US, they find that the real oil price is the most important factor determining real exchange rates in the long run. Camarero and Tamarit (2002) use panel cointegration techniques to investigate the relationship between real oil prices and the Spanish peseta's real exchange rate.
Koranchelian et al. (2005) finds that the long-run real exchange rate of Algeria is dependent on movements in relative productivity and real oil prices. Zalduendo (2006) using vector error correction model finds that increases in oil prices are associated with the appreciation pressures (and vice versa for price declines). There is also, however, a trend decline in the equilibrium rate that appears to be explained by depreciating pressures

arising from the sharp decline in productivity differentials recorded by the Venezuelan economy, against the backdrop of a marked increase in economic volatility. (Olomola and Adejumo, 2006) use quarterly data over the period 1970–2003 to examine the relationship between real oil price shock and real effective exchange rates, among other macro variables, for Nigeria. Applying the variance decomposition technique, based on a VAR model, they find that real oil prices lead to an appreciation of the real exchange rate.

(Chen and Chen, 2007) in a panel study of G7 countries showed that real oil prices may have been the dominant source of real exchange rate movements and there is a positive link between oil prices and real exchange rate. (Benassy-Quere et al., 2007) in the study of cointegration and causality between the real price of oil and the real price of the dollar over the 1974–2004 period found that, other things equal, a 10% rise in the oil price leads to a 4.3% appreciation of the dollar in real effective terms in the long run.

This paper adds to the existing literature on the determinants of real exchange rates in a number of different way as we focus on the three major oil exporting countries in Africa in terms of their current account surplus,Currently, not too many studies like this that run a thorough analysis of the real exchange rates in these three countries, using a consistent and coherent single-equation time-series approach and testing for the potential impact of real oil prices. From a methodological perspective, we build ourmeasures of the real effective exchange rates and productivity differentials against the 15 OECD main trading partners, thus controlling also for the so-called Balassa-Samuelson effect and interest rate differentials as a potential explanatory variable of the real exchange rate. Considering the fact that these three countries adopted different exchange rate regimes, we try to understand whether these arrangements may account for potential differences in the relationship between the real oil price and the real exchange rate. In particular, one would expect the relationship between oil prices and the real exchange rate to hold in countries where the nominal exchange rate is allowed to absorb potential exogenous oil shocks.

This paper is organized as follows. Section two describes the statistical characteristics of our data. Section exhibit econometric methodology and empirical results .We account for our findings in section 4. Section 5 concludes.

3.0 Data Description and Discussions:

The three economies here suggested, Nigeria, Algeria and Gabon are highly dependent on oil revenue.as earlier stated, oil exports accounts for around ninety percent of total exports in Gabon and contributes 45% to the nation's gross domestic product. In Nigeria, The oil and gas sector accounts for about 35 per cent of gross domestic product, and petroleum exports revenue accounts for about 70 per cent of total exports revenue For Algeria, the oil and gas sector contributes 35% to her gross domestic product and two-thirds of her total export. Since the world price hike in 2000, these three countries have benefited enormously and their revenue has been on the increase. For the purpose of this work, two main determinant of exchange rate; world oil price and interest rate determinant are discussed.

3.0.1 World Oil Price:

Following literature, Amano and van Norden(1998) motioned a two-sector link between oil price and exchange rate. Tradable (oil) and non-tradable (labour) sector, with two major assumption alongside the fact that there is constant return to scale in technology, that these sectors do not make economic profit and that input are mobile.it further states that the output price of the tradable sector is fixed internationally; hence the real exchange rate corresponds to the output price in the non-tradable sector. A rise in the oil price leads to a decrease in the labour (non-tradable) price so as to meet the competitiveness requirement of the tradable sector. If the non-tradable sector is more energy intensive than the tradable one, its output price rises and real exchange rate appreciates. The opposite applies if the non-tradable sector is less energy intensive than the tradable one. This is so that labour can be retained in the non-tradable part because excess increase reflecting the current oil price in the wages will make labour resort to leisure and work less; this however was refuted as Benassy-Quere et al., (2007) indicated that if oil there is a variation in oil price, it allows for direct effect on the interest rate depending on the oil-intensity of a country.

However, the likes ofKrugman (1983) and Golub (1983) believed that higher oil prices will transfer wealth from the oil importers to oil exporter thereby affecting the oil-exporting country positively.

3.0.2 Interest Rate Differentials:

Dornbusch, (1976) posited that the domestic money supply may grow faster the foreign money supply. This also may cause the nominal exchange rate to deviate from the position corresponding to purchasing power parity due to the slow response to price variables. As a result, the interest rate parity condition requires an overshooting exchange rate. An overshooting exchange rate together with a slow adjustment of price levels generates a change in the real exchange rate. His theory suggests that money could have a temporary influence rather than long-term impact on the real exchange rate. When prices catch up after the disturbance occurs, the real exchange rate will move back to the original position.

3.1 Theoretical Inference:
Meese and Rogoff (1988) examined the co-movements of major currency real exchange rates and long-term real interest rates over the modern (post-March 1973) flexible exchange rate experience. The real exchange rate, *qt*, can be defined as:
qt≡et–pt+pt* (1)where *e*t is logarithm of nominal exchange rate (domestic currency per foreign currency unit) and *p*t and *p*t* are the logarithms of domestic and foreign prices. Three assumptions are made: first, that when a shock occurs, the real exchange rate returns to its equilibrium value at a constant rate; second, that the long-run real exchange rate, , is a non-stationary variable; finally, that uncovered real interest rate parity (UIP) is fulfilled:
*E*t(*q*t+k–*q*t)=*R*t– *R*t* (2)
where *R*t* and *R*t are respectively, the real foreign and domestic interest rates for an asset of maturity *k*. Combining the three assumptions above, the real exchange rate can be expressed in the following form:
*q*t=δ(*R*t–*R*t*)+t (3) where δ is a positive parameter larger than unity. This leaves relatively open the question of which are the determinants of t that is non-stationary variable. Equation (3) is the second relationship investigated in this paper and represents a typical model of the relationship between the real interest rate differential and the real exchange rate explored in the literature. When shocks are primarily real this relationship is likely to outperform the relationship between nominal exchange rates and real interest rate differentials that can also be derived using international parity conditions (see Meese and Rogoff (1988)).
Consequently, for the purpose of this paper, the real exchange rate (REER) is given as a function of real oil price(ROLP) and interest rate differential (IRD).

REER=F(ROLP,IRD) (4)

3.2 Econometric Model and Estimation:
From the aforementioned, this paper uses quarterly data of oil price, exchange rate and interest rate for panel of Nigeria, Algeria and Gabon from 1980Q1 to November 2012Q4. Source of data is WDI (World Development Indicator) and IFS (International Financial Statistics). Real exchange rates are constructed by using domestic price level and price level in a foreign country. Real exchange rate is equal to Nominal Exchange Rate * (Foreign Price Level / Domestic Price Level). Real oil price are defined as the price of Dubai crude oil expressed in US dollars, deflated by domestic consumer price index. Real oil price and real exchange rate are expressed in natural logarithm form. Real interest rate differentials (IRD) is calculated asIRDit= *rit– rt** , where *rit*is the real interest rate of country *i* and *rt** is the real foreign interest rate. Real interest rate is derived using Fisher equation. The real interest rate solved from the Fisher equation is (1 + Interest) divided by (1+Inflation) -1. The model to estimate is given as:
qit=αi+β1irdit-β2iroilt(5)
Where the exchange rate (*qit*) is defined as the cost of a unit of foreign currency in terms of the domestic currency, *drrit*is the real interest rate differential and *roilpit*is the real price of oil. According to the theoretical model, an increase in the real interest rate differential would appreciate the currency. The sign corresponding to the real price of oil would be negative for oil exporting countries. This is because an increase in the real price of oil will depreciate the oil importing currencies relative to oil exporters. Where β1i < 0 and β2i < 0.
However, before this estimation can be done,there is a need to determine the integration of the three variables in this panel. An integrated series needs to be differenced in order to achieve stationarity. A panel series Yit, that requires no such differencing to obtain stationarity is denoted as Yit ¡«I(0). Therefore, an integrated series such as Yit ¡«I(1) is said to grow at a constant rate while Yit ¡«I(0) series appear to be trendless. Thus, if two series Yit and Xit are integrated of different order, say Yit ¡«I(0) and Xit ¡«I(1) respectively, then they must be drifting apart over time. Therefore, a regression of Yit on Xit would encounter a spurious regression problem, as the residual would also be I(1) which violates the underlying assumptions of ordinary least squares (OLS). Thus, it is important to determine that the series of interest have the same order of integration before proceeding into further estimation.
After establishing the order of integration of the data, the paper would use panel cointegration approaches to test for a long run equilibrium relationship among variables. If two series Yit and Xit are both I(1) then it is normally the case that a linear combination between the two will also be I(1) so that a regression of Yit on Xit would produce spurious results. This is because the residual is also I(1), which violates the assumptions of OLS. However, in a special case, a linear combination of two I(1) variables will result in a variable (residual) which is I(0). (Granger, 1981) has called such variables cointegrated. As shown by (Engle and Granger, 1987), there must be a vector error correction representation governing the comovements of these series over time. This leads to the intuitive interpretation of a cointegrated system as one that represents long-run steady state equilibrium.
Generally, if two or more variables are cointegrated, there is a long-term equilibrium relationship between them. To investigate the long-run relationship between the variables under study, the paper will adopt panel estimation method instead of standard OLS regression. With non-stationary variables, an OLS regression suffers from serial correlation. Moreover, since the cointegration literature does not assume exogenous regressors, estimation must account for potential endogenous feedback between X and Y (Funk, 2001). The advantage of panel estimators

over standard time-series regressions is that each estimator is super-consistent. Asymptotically, the OLS estimator is normal with a nonzero mean, while panel estimators such as the PMG estimator proposed by Pesaran et al., (1999) are normal with zero means irrespective of whether the underlying regressors are I(1) or I(0).

The methods applied to the estimation of the real exchange rate model are based on the combination of panel techniques and cointegration tests. The first step to take, as in the time series context, is to analyze the order of integration of the variables, as a pre-requisite. The paper employs several panel data unit root tests in order to exploit the extra power in the cross-sectional dimension of the data. Specifically, the paper utilizes the panel unit root tests proposed by (Levin et al., 2002), (Breitung, 2000), (Im et al., 2003), (G. S. Maddala, 1999) (1999) and (Hadri, 2000). Levin et al., (2002), Breitung (2000), and Hadri (2000) tests all assume that there is a common unit root process so that ρi is identical across cross-sections. The first two tests employ a null hypothesis of a unit root while the Hadri (2000) test uses a null of no unit root. Levin et al. (2002) and Breitung (2000) consider panel versions of the Augmented Dickey–Fuller (ADF) unit root test (with and without a trend). These tests restrict α to be identical across cross-sectional units, but allow the lag order for the first difference terms to vary across cross-sectional units, which in this study are countries.

$$\Delta y_{it} = k_i + \propto y_{it-1} + \sum_{j=1}^{x} \varphi \ ij \Delta y_{it-j} + \in_{it} \qquad (6)$$

$$\Delta y_{it} = k_i + \propto y_{it-1} + \beta t \sum_{j=1}^{x} \varphi \ ij \Delta y_{it-j} + \in_{it} \qquad (7)$$

The subscript i=1, c,N indexes the countries. Equations (6) and (7) are estimated using pooled ordinary least squares (OLS). Levin et al. (2002) tabulate critical values for ta by performing Monte Carlo simulations for various combinations of N and T commonly employed in applied work. The null and the alternate hypotheses are: H0: $f_{¿}$=0 and H1: $f_{¿}$<0. Under the null hypothesis there is a unit root, while under the alternative hypothesis, there is no unit root. The difference between the Levin et al. (2002) test and the Breitung (2000) test is that while the former requires bias correction factors to correct for cross-sectionally heterogeneous variances to ensure efficient pooled OLS estimation, the Breitung (2000) test achieves the same result by appropriate variable transformations (Narayan et al., 2008).

One of the drawbacks of the Levin et al. (2002) and Breitung (2000) tests is that in Equations (6) and (7) $\propto$ is restricted to be identical across countries under both the null and alternative hypotheses. The t-bar test proposed by Im et al. (2003) has the advantage over the Levin et al. (2002) and Breitung (2000) tests that it does not assume that all countries converge towards the equilibrium value at the same speed under the alternative hypothesis and thus is less restrictive. (Karlsson and Löthgren, 2000) perform Monte Carlo simulations that show that in most cases the Im et al. (2003) test is superior to the Levin et al. (2002) test. There are two stages in constructing the t-bar test statistic. The first is to calculate the average of the individual ADF t-statistics for each of the countries in the sample. The second is to calculate the standardized t-bar statistic according to the following formula:

t– bar = root of **N** (tá – êt)/root of v_t (8)

where N is the size of the panel, tα is the average of the individual ADF t-statistics for each of the countries with and without a trend and κt and νt are, respectively, estimates of the mean and variance of each tαi. Im et al. (2003) provide Monte Carlo simulations of κt and νt and tabulate exact critical values for various combinations of N and T. A potential problem with the t-bar test is that when there is cross-sectional dependence in the disturbances, the test is no longer applicable. However Im et al. (2003) suggest that in the presence of cross-sectional dependence, the data can be adjusted by demeaning and that the standardized demeaned t-bar statistic converges to the standard normal in the limit.

Maddala and Wu (1999) criticize the Im et al. (2003) test such that cross correlations are unlikely to take the simple form proposed by Im et al. (2003) in many real world applications that can be effectively eliminated by demeaning the data. Maddala and Wu (1999) propose an alternative approach to panel unit root tests using Fisher's (1932) results to derive tests that combine the p-values from individual unit root tests. The test is non-parametric and has a chi-square distribution with 2N degrees of freedom, where N is the number of cross-sectional units or countries. Using the additive property of the chi-squared variable, the following test statistic can be derived:

$$\lambda = -2 \sum_{i=1}^{n} log_e \pi_i \qquad (9)$$

Here, πi is the p-value of the test statistic for unit i. An important advantage of this test is that it can be used regardless of whether the null is one of integration or stationarity. The paper also implemented the panel stationarity test suggested by Hadri (2000). The Hadri (2000) panel unit root test is similar to the (Kwiatkowski et al., 1992) unit root test, and has a null hypothesis of no unit root in any of the series in the panel. Like the Kwiatkowski et al. (1992) test, The Hadri (2000) test is based on the residuals from the individual OLS regressions from the following regression model:

yit=πi+θit+µit (10)

Given the residuals û from the individual regressions, the LM statistic is:

$$LM = \frac{1}{N}(\sum_{i=1}^{N}\sum_{t} \quad S_i(t)^2/T^2/\,f_o) \qquad (11)$$

Where *Sit* are the cumulative sum of the residuals,

$$s_i(t) = \sum_{i=1}^{t} \quad U_{it} \qquad (12)$$

f is the average of the individual estimators of the residual spectrum at frequency zero

$$f = \frac{\sum_{t=1}^{N} \quad f_{io}}{N} \qquad (13)$$

Hadri (2000) shows that under mild assumptions,

$$\delta = \frac{\sqrt{N}(LM-\xi)}{\varphi} \qquad (14)$$

Where ξ = 1/6 and ξ = 1/45 and φ=1/45, if the model only includes constants (is set to 0 for all), and ξ = 1/15 and φ = 11/6300 , otherwise. It is worth noting that simulation evidence suggests that in various settings (for example, small T), Hadri's panel unit root test experiences significant size distortion in the presence of autocorrelation when there is no unit root. In particular, the Hadri (2000) test appears to over-reject the null of stationarity, and may yield results that directly contradict those obtained using alternative test statistics (see (Hlouskova and Wagner, 2006) for discussion and details).

3.3 Panel Unit Root Results:

The table below shows the panel unit root test results, there are three different null hypotheses for the panel unit root tests. The first two are the Breitung (2000) and Levin et al. (2002) tests where the null hypothesis is the unit root (with the assumption that the cross-sectional units share a common unit root process). The second group includes two tests (Im et al. (2003), and Maddala and Wu (1999) Fisher type test with null of unit root assuming that the cross-sectional units have individual unit root process. The last test is the Hadri (2000) test, where the Z-stat has a null hypothesis of no unit root (but assumes a common unit root process for all cross-sectional units). All test results are based on the inclusion of an intercept and trend.

Table 1.1 Panel Unit Root Test Results In Levels

	Null Hypothesis	Exchange rate	Oil price	Interest rate differential
Levin, Lin and Chu	Unit Root	1.24819 (0.8940)	-0.17553 (0.4303)	-4.99959 (0.0000)
Breitung t-stat	Unit Root	-0.39706 (0.3457)	1.65883 (0.9514)	-4.32091 (0.0000)
Im, Pesaran& Shin	Unit Root	1.85981 (0.9685)	1.97221 (0.9757)	-3.51340 (0.0002)
ADF-Fisher Chi-square	Unit Root	0.84330 (0.99809)	1.26674 (0.9735)	22.7171 (0.0009)
Hadri Z-stat	Stationary	6.64177 (0.0000)	11.3038 (0.0000)	2.99205 (0.0014)

Table 1.2 Panel Unit Root Results In First Difference.

	Null Hypothesis	Exchange rate	Oil price	Interest rate differential
Levin, Lin and Chu	Unit Root	-22.4952 (0.0000)	-24.7592 (0.0000)	-10.1320 (0.0000)
Breitung t-stat	Unit Root	-19.3701 (0.0000)	-19.0605 (0.0000)	-7.95752 (0.0000)
Im, Pesaran& Shin	Unit Root	-20.8219 (0.0000)	-21.8393 (0.0000)	-12.5450 (0.0000)
ADF-Fisher Chi-square	Unit Root	212.387 (0.0000)	221.025 (0.0000)	123.049 (0.0000)
Hadri Z-stat	Stationary	-0.02223 (0.5089)	-0.12596 (0.5501)	2.79835 (0.0026)

Note: (a) An intercept and trend are included in the test equation. The lag length was selected by using the Modified Akaike Information Criteria.

(b) Probability written in parenthesis.

It is clear that real oil price, interest rate differentials and real exchange rates are $I(1)$ series for panel of three countries. For real oil price, each of the five tests suggest stationarity at first difference at 1% level of significance. Same goes for real exchange rate, without anyexception. For real interest rate differential, Hadri's Z-stat rejects null of stationarity and Levin et. al. (2002) test rejects null of non-stationarity at 1% significance level in every case. To sum up, the results indicates that there is stationarity in first differences and each of the three variables can be regarded as $I(1)$. In what follows, the paper will proceed on the assumption that all variables are $I(1)$ and differenced variables are $I(0)$. In such a case as this, Cointegration methods would be appropriate and therefore applied. This is explained below.

3.4 Panel Cointegration Tests

This test is carried out to check for the presence of cointegration which is a check for long run relationship between exchange rate, real oil price and real interest rate differential variables. The paper utilise panel cointegration tests due to Pedroni (1998), Kao (1999) and Maddala and Wu (1999). The tests proposed in (Pedroni, 1998) are residual-based tests which allow for heterogeneity among individual members of the panel, including heterogeneity in both the long-run cointegrating vectors and in the dynamics. Two classes of statistics are considered in the context of the Pedroni (1998) test. The panel tests are based on the within dimension approach (i.e. panel cointegration statistics) which includes four statistics: panel v-statistic, panel ñ-statistic, panel PP-statistic, and panel ADF-statistic. These statistics essentially pool the autoregressive coefficients across different countries for the unit root tests on the estimated residuals. These statistics take into account common time factors and heterogeneity across countries. The group tests are based on the between dimension approach (i.e. group mean panel cointegration statistics) which includes three statistics: group ñ-statistic, group PP-statistic, and group ADF-statistic. These statistics are based on averages of the individual autoregressive coefficients associated with the unit root tests of the residuals for each country in the panel. All seven tests are distributed asymptotically as standard normal. Of the seven tests, the panel v-statistic is a one-sided test where large positive values reject the null hypothesis of no cointegration whereas large negative values for the remaining test statistics reject the null hypothesis of no cointegration.

The (Kao, 1999) test follows the same basic approach as the Pedroni (1998) tests, but specifies cross-section specific intercepts and homogeneous coefficients on the first-stage regressors. In the null hypothesis, the residuals are nonstationary (i.e., there is no cointegration). In the alternative hypothesis, the residuals are stationary (i.e., there is a cointegrating relationship among the variables). The third test is the Johansen-type panel cointegration test developed by Maddala and Wu (1999). The test uses Fisher's result to propose an alternative approach to testing for cointegration in panel data by combining tests from individual cross-sections to obtain at test statistic for the full panel. The Maddala and Wu (1999) test results are based on p-values for

Johansen's cointegration trace test and maximum eigenvalue test. Evidence of cointegration between real exchange rate and real oil price using the Maddala and Wu (1999) test is obtained if the null hypothesis of none (r = 0) cointegration variables is rejected and the null of at most 1 (r ≤ 1) cointegrating variables is accepted, suggesting the direction of causality is running from real oil price to real exchange rate. In other word, the paper would confirm the existence of a unique cointegration vector for the estimated model.

3.5 Panel Cointegration Tests Results

Table 2.0Pedroni(1998) panel cointegration test Results

Within Dimension		Between Dimension	
Test statistics	Value	Test statistics	Value
Panel v-Statistic	-0.193561 (0.5767)	Group rho-Statistic	0.976026 (0.8355)
Panel rho-Statistic	-0.003282 (0.4987)	Group PP-Statistic	0.628036 (0.7350)
Panel PP-Statistic	-0.131326 (0.4478)	Group ADF-Statistic	0.691182 (0.7553)
Panel ADF-Statistic	-0.058448 (0.4767)		

Table 2.1 Kao (1999) Residual CointegrationTests Result

Null hypothesis: No cointegration	Statistics	Probability
Panel	-0.499214	0.3088

Table 2.2 Maddala& Wu (1999) Fisher Panel Cointegration Test Results

Hypothesized No. of CE(s)	Fisher Stat. (from trace test)	Probability	Fisher Stat.* (from max-eigen test)	Probability
None	29.41	0.0001	25.76	0.0002
Almost 1	10.81	0.0943	6.080	0.4143
Almost 2	16.94	0.0095	16.94	0.0095

Table 2.0 to 2.2 from above reports three types of cointegration test results, the panel tests of Pedroni (1998) indicate no support for the hypothesis that real oil prices and real interest rate differential are cointegrated with real exchange rate for panel of the three countries. Same is the case with kao(1999). However, evidence of cointegrating relationship between the variables are obtained from Maddala& Wu (1999) tests. The null hypothesis of no cointegrating relationship is rejected at 1% level for panel.in lieu of this fact, the paper therefore continues with econometric technique which takes into account this long-run relationship between the variables

4.0 Long run estimate

In the third step, having found that a cointegrating relationship holds among real exchange rate, real oil price and real interest rate differential for the panel of eight countries and for each respective country group, the paper proceeds with the estimation of the long-run elasticities on the impact of real oil price and real interest rate differential on real exchange rate. The estimation of real exchange rate equilibrium model is based on pooled cross-country time series data. The main advantage of panel data for the analysis of real exchange rate equations is that the country-specific effects can be controlled for, for example by using dynamic fixed effect (DFE) estimator. However, such approach generally imposes homogeneity of all slope coefficients, allowing only the intercepts to vary across countries. (Pesaran and Smith, 1995) suggest that, under slope heterogeneity, this estimate is affected by a potentially serious heterogeneity bias, especially in small country samples.

Conversely, the mean group (MG) approach due to Pesaran and Smith (1995) allows all slope coefficients and error variances to differ across countries, having considerable heterogeneity. The MG approach applies an OLS method to estimate a separate regression for each country to obtain individual slope coefficients, and then averages the country-specific coefficients to derive a long-run parameter for the panel. For large T (the number of time periods) and N (the number of units), the MG estimator is consistent. With sufficiently high lag order, the MG estimates of long-run parameters are super-consistent even if the regressors are nonstationary (Pesaran et al., 1999). However, for small samples or short time series dimensions, the MG estimator is likely to be inefficient (Hsiao et al., 1999). For small T, the MG estimates of the coefficients for the speeds of adjustment are subject to a lagged dependent variable bias (Pesaran et al., 1999b).

Unlike the MG approach, which imposes no restriction on slope coefficients, the pooled mean group (PMG) estimators due to Pesaran et al. (1999a) allow short-run coefficients, speed of adjustment and error variances to differ across countries, but impose homogeneity only on long-run coefficients. This estimator is especially suited for panels with large T and N. It does not impose homogeneity of slopes in the short-run and it allows for dynamics. Therefore, under the null hypothesis of long-run homogeneity across coefficients, the paper estimates the long-run elasticities of the impact of real oil prices and real interest rate differential on real exchange rate equation on monthly data for 8 countries from 1980 to 2008 using the PMG procedure. In practice, the PMG procedure involves first estimating autoregressive distributed lag (ARDL) models separately for each country i.

$$q_{it} = \mu_{it} + \xi_{t,1} lnoilp_{it} + \xi_{t,2} lnoilp_{i,t-1} + \varphi_{i,1} ird_{it} + \varphi_{i,2} ird_{it-1} + \theta_i q_{it-1} + \varepsilon_t \quad (16)$$

From the above equation, subscript I represent individual country cross-sectional unit while t is the time period, However, the corresponding error correction equation is written below

$$\Delta q_{it} = \varphi_2 (q_{it-2} - a_{i1} - a_{i2} lnoilp_{it} - a_{i3} ird_{it}) - \xi_{t,1} \Delta lnoilp_{it} - \psi_{t,1} \Delta ird_{it} + \upsilon_{it} \quad (17)$$

Where $a_{i1} = (\frac{\mu_i}{1} - \theta_i)$. $a_{i2} = (\xi_{i,1} + {}^{\xi_{t,2}}/_{1} - \theta_i)$ and $a_{i3} = (\psi_{i,1} + \frac{\psi_{t,2}}{1} - \theta_i$.

From the above Equation is the coefficient that measures the speed of adjustment to short-run disequilibrium, $a_{i2} lnoilp_{it}$ and $a_{i3} ird_{it}$ are the long run coefficients of real oil price and real interest rate differential respectively while $\xi_{t,1} \Delta lnoilp_{it}$ and $\psi_{t,1} \Delta ird_{it}$ are the short run coefficients for real oil price and real interest rate differential respectively. For purpose of robustness check, the paper also employs pooled mean group (PMG) as the methodology to estimates the long run effects. The long-run slope homogeneity hypothesis of PMG is tested via the Hausman test. Under the null hypothesis, PMG estimators are consistent and more efficient than MG estimators and DFE, which impose no constraint on the regression (Pesaran et al., 1999).

4.1 Estimation Results:

Pool mean group

Dependant Variable: LNREER	Without time trend	With time trend
Short run		
Coefficient	-0.002684 (0.1846)	-0.003358 (0.1022)
lnoilp	0.083306 (0.0148)	-0.006353 (0.8444)
ird	0.002024 (0.0000)	0.001801 (0.0000)
Long run		
coefficient	-0.002727 (0.1850)	-0.003382 (0.1030)
lnoilp	0.083179 (0.0159)	-0.005970 (0.8554)
ird	0.002024 (0.0000)	0.001801 (0.0000)
Husman	0.723122 (0.0023)	
No of countries: 3 Observation: 396		

From the table above, the impact of real oil price and real interest rate differential on real exchange rate for net oil exporting countries. The long run restriction imposed by PMG estimators cannot be rejected at 1% level by the Hausman test statistics for both specifications. The PMG estimates however finds no evidence to suggest that real oil price has negative effect on real exchange ratesboth in the short and long run without time trend but finds strong evidence for real interest differential at 1% level of significance.. The coefficient on real oil price for PMG estimates are negatively signed, although not significant in the long run, a result that is consistent with previous studies based on oil exporting countries. Perhaps the lack of evidence to suggest that real oil price has positive effect on real exchange rate is due to the choice of sample countries included in the estimation. Of three net oil exporters in the sample, Gabon registered a positive correlation between real oil price and real exchange rate (when it should had been negative). Pooling these countries together may yield inconsistent slope coefficients among individual sample countries hence resulting in insignificant long run estimation results.

Taking into account the whole set of regression results, this analysis on quarterly data clearly shows a significant effects of real oil price and real interest rate differential on real exchange rate when using the PMG approach. This is true mainly for panel of three countries. The findings in general suggest that higher real oil price would results in depreciation in real exchange rate for net oil exporting countries. On the impacts of real interest rate differential on real exchange rates, the PMG estimates provide evidence for positive long run relationship between the variables.

5.0 Summary and Conclusion

This research based on establishing the existence of a link between oil price and exchange rate for three oil exporting countries applying very recent tests of unit root and cointegration in a panel data based on the simple model of messe and rogoff (1988) decipher the evidence for any long run relationship among real exchange rate, real interest rate differential and real oil price for the period 1980q1 to 2012q4. The use of these methods, quite recent in the applied literature, avoids the problems found in panel data analysis when the variables are non-stationary, and adds the cross-country dimension to the traditional time series analysis. The inclusion of the real interest rate differential and real oil price as the determinant of the equilibrium real exchange rate seems to provide a reasonable model to explain the behaviour of the real exchange rate among net oil importing countries in particular.

However, the paper found that all the three variables are stationary at first difference at 1% significance. This paper also proved the existence of a long run relationship between the three series. While the Pedroni (1999) andkao (1999) test failed to find evidence of cointegration, Maddala and Wu (1999) test provide significant evidence of cointegration among the variables for all three of countries.

Conclusively, to investigate the impacts of real oil price on real exchange rate, the paper conducted a dynamic panel data study allowing for considerable heterogeneity across countries for three countries over 1980-2012 using the pooled mean group (PMG) methodology which allows for heterogeneous dynamic adjustments towards a common long-run equilibrium. This research in general provides strong evidence in support of a significant positive impact of real oil price and real interest rate differential on real exchange rate, indicating any future oil price shocks would cause real depreciation of exchange rate in the long run especially among net oil exporting countries in the short run. The paper however did not find evidence to suggest that higher real oil prices lead to real appreciation of exchange rate among net oil exporting countries. This is nevertheless not surprising because previous literatures which attempted to link the effect of real oil price on real exchange rate for oil exporting countries were based on OPEC countries where oil accounts for at least three-quarters of total export earnings. Notwithstanding, strong evidence is obtained to link between real interest rate differential and real exchange rate for individual country. The sign of oil price coefficient is negative in the long run and is in tandem with the theory.

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