Oil Price and Real Exchange Rate Appreciation: Is there Dutch Disease in Nigeria?

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Abstract

Nigeria has witnessed oil windfalls for decades, however, it was presumed that revenues from oil would assist Nigeria to mark its transition from underdevelopment to industrial development. Regrettably, the high dependence on oil sector has led to deindustrialization and exchange rate appreciation. This paper studies the dynamic relationship between oil price and real exchange rate for Nigeria using annual time series data over the sample period 1970 to 2018. Our analysis is based on Autoregressive distributed lag bound testing to cointegration approach. Following the integration and cointegration tests, the results indicate that all the variables attained long run cointegrating relationship. The results also indicate that oil price, government expenditure and inflation contributes to real exchange rate appreciation, thus, confirming the existence of Dutch disease. Essentially, this is a clear evidence that Dutch disease weakens the competitiveness of the lagging sectors in Nigeria. Also, the paper identifies unidirectional causality from oil price to real exchange rate, and bidirectional causality between oil price and government expenditure, real exchange rate and inflation. This means that the oil price and inflation encourages real exchange rate appreciation. Consequently, the policy implication is to diversify the lagging sectors by investing in human resource development, basic infrastructure and tax concessions will raise the competitiveness level of the lagging sectors.

Keywords: Real effective exchange rate, Oil price, Government expenditure, inflation

JEL: F30; O13; H5

1. Introduction

The term Dutch disease is derived from the shrinking of the manufacturing sector in the Netherland after the discovery of natural gas (Economist, 1977). Since then, Dutch disease has been used as a tool for clarifying the natural resource curse. In particular, the Dutch disease describes a situation where the windfall of natural resource might lead to appreciation of real exchange rate and the competitiveness of the lagging sectors being squeezed out (Beine et al., 2012; Adamu and Rajah, 2017). However, deindustrialisation, which contributes to the reallocation of wages and labour away from the lagging sectors, is the most frequent observed consequence of this scenario. The crowding out of these sectors not only diversifies the oil and gas sector, but also extends the non-traded goods sector. This phenomenon leads to inflation and high government spending, and associated with the real exchange rate appreciation occasioned by the two events. Firstly, when commodity boom is cyclic with its negative effects and turns to commodity bust. Secondly, it can be a persistent phenomenon, especially when resource rich countries often establish a social structure in which corrupt political elites are supported by physically control of natural resources (Asekunowo et al., 2012; Adamu, 2016).

Why Nigeria is suitable case for this study? Prior to independence and afterwards, agriculture was the key pillar of the Nigerian economy, contributing morethan 50 percent of GDP and employing 72 percent of the labour force (see, Olusi and Olagunju, 2005). Furthemore, morethan 60 percent of the Nigeria's foreign exchange earnings were generated from agricultural exports. The oil boom of the 1970s, however, had a significant impact on the Nigerian economy. For example, the country has earned more than US$400billion from oil and gas-related fiscal revenues from 1970 to date, indicating that oil exports have dominated agricultural exports contributing more than 90% of government foreign exchange earnings and about 54% of the country's GDP, 10% and 96% of employment and export earnings per year (Olusi and Olagunju, 2005; Budina and van Wijinbergen, 2007; Adamu, 2017). These have allowed massive investment plans and increasingly growing government expenditure to diversify the economy (Adamu and Rajah, 2017; Adamu, 2016). This suggests that natural resource boom causes deindustrilisation, which in turn leads to real appreciation of exchange rate and shrinks the competativeness of the lagging sectors.

The aim of this study is to investigate whether or not the Dutch disease exists in Nigeria. Specifically, this study uses the real exchange rate as proxy for Dutch disease to empirically analyse whether the Dutch disease exists in Nigeria, which is rich in oil. Despite the vast literature on Dutch disease, this paper seeks to extends the existing literature in threefold. First, we observed that relatively few attempt is made to assess the consequences of the oil boom that might cause Dutch disease in Nigeria. For instance, Adenikinju *et al.*, (2002) and Olusi and Olagunju (2005) have explored the issue and their empirical findings are generally inconclusive. Second, we consider structural breaks in fortifying the order of integration since conventional unit root test might be one sided if the variables are encountered with a structural breaks and is overlooked (Perron, 1989). Third, examining the direction of causality plays a significant role in planning for an appropriate policy, however, to our knowledge, this study is among the few that evaluate the causal relationship among the exogeneous variables under examination. The results would provide an avenue to fine tune policies to minimize the effect of Dutch disease in Nigeria.

The rest of the paper is structured as follows. Section 2 includes an overview of the literature. The methodology and data are discussed in section 3 while results and discussion are provided in section 4. Finally, the paper ends in section 5 with a conclusion and policy implications.

2. Literature Review

Constraints of long-term economic growth does not stem from the external restraints or capacity constraint, but from an appreciation of real exchange rate occasioning from Dutch disease (Marconi *et al.,* 2014). Dutch disease is a phenomenon that simply denotes the persistent appreciation of real exchange rates induced by exports of commodities making use of low cost and plentiful natural resources, and generating substantial export revenues for the economy (Bresser-Pereira, 2008, Adamu, Bawa and Tukur, 2020). The central theory is that the tradeable sector booms while the lagging sector shrink. Specifically, the theory informs us that significant volume of commodity exports will give rise to an appreciation of the real exchange rate and consequently weaken the lagging sectors exposed to foreign competition (Corden, 1984; Corden and Neary, 1982). Sachs and Warner (1995) point that since both the prices of booming and lagging sectors of export and import manufactures are determined by world market prices and the exchange rate, therefore, resource-rich countries tend to be vulnerable to 'Dutch Disease' because a considerable number of these countries are mainly dependent on export of natural resources as the major source of income. Accordingly, these countries domestic currencies appreciate in value compared to other currencies, making their goods expensive and thus less competitive. Hence, depending on wealth of natural resource, it hampers other export-led-economic activities, particularly manufacturing, and such economies are prone to stagnation structurally and have little, if any, diversification (Van Wijinbergen, 1985).

Even though a number of empirical studies have supported the positive effect of oil price on real effective exchange rate of the oil-rich countries thereby establishing the existence of Dutch disease, On the other hand, there are number of empirical evidences that have conflicting view on similar relationship. For example, Adenikinju *et al.*, (2002) assess manufacturing competitiveness in Africa using Cameroon, Cote D'Ivoire, Nigeria and Senegal. They established that Nigeria, Cameroon and Code D'Ivoire are affected by "Dutch disease" and Senegal had the least symptoms of terms of trade variations. Olusi and Olagunju (2005) investigates the validity or otherwise of the Dutch disease in Nigeria. The findings show the presence of Dutch disease in Nigeria. For Azerbaijan economy, Gahramanov and Fan (2005) report that the prevalence of Dutch disease was less vulnerable. Though, government ought to provide measures against future risk.

Furthermore, Kutan and Wyzan (2005) extends the Balassa-Samuelson model to investigate the possibility or otherwise of Dutch disease effects in Kazakhstan. Results indicate oil price, inflation and productivity have a significant impact on the changes in the real exchange rate, which is a manifestation of the Dutch disease. Olomola and Adejumo (2007) explores the nexus between real oil price and real exchange rate and conclude that real oil prices affect the appreciation of the real exchange rate.

Egert and Leonard (2008) study the Kazakhstan economy to establish symptoms or the existence of Dutch disease. Results indicate that higher oil prices may be related to an appreciation of the real exchange rate. Javaid and Riazuddin (2009) investigates Dutch disease hypothesis in some selected South East Asian countries. The study conclude that real exchange rate appreciation is caused by increase in foreign inflows and contraction in the tradable and expansion in the non-tradable sector.

In a similar study, Pegg (2010) reports that Botswana has nothing to do with either diamond export earnings or real exchange appreciation issues, but symtoms of Dutch disease may appear in the near future. Jbir and Zouari-Ghorbel (2011) investigates the oil price and Dutch disease nexus for Algeria. The empirical results show that higher inflation as a result of oil price shocks led to excessive spending was the cause of Dutch disease.

Algieri (2011) studies if the the Russian economy exhibits the symptoms of the Dutch Disease over the transition period commenced in the early 1990s. The findings indicate oil price shock leads to a real exchange rate appreciation, a rise in output growth and a decline in local manufacturing output. Otaha (2012) analyses the causes of Dutch disease in Nigeria and concludes that increasing oil windfalls contributes to an increase in exchange rates and stirs the balance of payments problem. Sala-i-Martin and Subramanian (2013) examines the Dutch disease hypothesis in Nigeria and discover that oil price shocks does not determines real exchange rate appreciation.

Dulger *et al.*, (2013) examines the Dutch disease phenomena for Russia after the de-industrialization in the Post-Soviet union. Finding reveals an evidence of Dutch disease in the Russia. Ito (2017) explores whether or not the Dutch disease is present in Russia. The results suggest that manufacturing output positively related to the oil price and increases marginally, even though the real effective exchange rate appreciates. The FDI inflows add to manufacturing output but not significant while increasing government spending crowds out the manufacturing sector.

3. Methodology

3.1 Model and Data

To investigate the existence or symptoms of Dutch Disease for Nigeria, we formulate a Dutch Disease model by incorporating macroeconomic variables that are standard in the literature to explain the relationship. The Dutch disease functional type is written as:

$$rer_t = f(olp_t, gex_t, ifn_t) \tag{1}$$

Equation (1) is transformed into an empirical model for the base line long run equation as:

$$\ln rer_t = \alpha_0 + \alpha_1 \ln olp_t + \alpha_2 \ln gex_t + \alpha_3 ifn_t + \varepsilon_t \tag{2}$$

where *rer* = real exchange rate (proxy for Dutch disease), *olp* = oil price; *gex* = government expenditure; *ifn* = inflation rates; are log of real exchange rate, log of oil price, log of government expenditure and inflation. $a_0 - a_3$ = parameters to be estimated; t = time period; ln = natural log and ε = stochastic term.

According to Dutch desease theory resource boom should have a positive effect on real exchange rate (Corden, 1984; Valeriy and Anna, 2015), therefore, the coefficient of oil price (*olp*) should be positive. Beside oil price, there are variables that have been established in the literature. For example, resource boom such as influx of oil revenues encourages high government spending in a way that would be difficult to reverse when it turns to burst (Adamu, 2019). Therefore, any increase in government spending may cause appreciation in real exchange rate, thus, positive sign is expected. It is expected that increase in government spending, triggered by higher oil prices may leads to inflation (*ifn*) (Rosenberg and Saavalainen, 1998; Engel, 2002; Jbir and Zouari-Ghorbel, 2011). Therefore, resource abundant countries exprecienced inflation after post-resource windfalls and puts pressure on the real exchange rate, positive sign is expected.

The study employs annual time series data for the sample period 1970 to 2018. As for the Dutch disease, there is no clear index on non-tradables and tradable goods, we use real effective exchange rate as proxy for Dutch disease, government expenditure as percentage of GDP and inflation are collected from the Central Bank of Nigeria and the World Development Indicators, World Bank while data on oil price is obtained from the Organisation Petroleum Exporting Countries (OPEC).

3.2 Integration and ARDL Cointegration tests

As a pre-condition in dealing with the time series data, and to avoid spurious regression, the properties of the candidate variables must be examined (Engle and Granger, 1987). This study employs two traditional unit root tests for this analysis - Augmented Dickey-Fuller (1981) and Phillips and Perron (1988) unit root tests. However, most of the time series data have structural break owing to economic reform and irregular shocks, and ADF and PP unit root tests may not be appropriate when the series contains structural break. To overcome this problem, we employ unit root test proposed by Zivot and Andrew (1992) to account for structural breaks. This unit root test provides information on single break point at particular period endogenously from the series. Furthermore, it does not provide test results that biased to conventional unit roots. Zivot and Andrew use three models. Model A allows for change in trend and intercept, Model B allows for change in trend and slope, while Model C allows change in trend, intercept and slope respectively. The models are in the following form:

Model A: $$\Delta y_t = \psi + \gamma_1 y_{t-1} + \alpha t + \lambda_1 DU_t(\delta) + \sum_{j=1}^{p} \phi_j \Delta y_{t-j} + \varepsilon_t \quad (3)$$

Model B: $$\Delta y_t = \psi + \gamma_1 y_{t-1} + \alpha t + \psi_1 DT_t(\delta) + \sum_{j=1}^{p} \phi_j \Delta y_{t-j} + \varepsilon_t \quad (4)$$

Model C: $$\Delta y_t = \psi + \gamma_1 y_{t-1} + \alpha t + \lambda_1 DU_t(\delta) + \psi_1 DT_t(\delta) + \sum_{j=1}^{p} \phi_j \Delta y_{t-j} + \varepsilon_t \quad (5)$$

Where Δ is the first difference operator, y_t is the time series to be tested and t = 1, 2, 3, 4, . . ., T. DU_t stand for the Dummy variable for a mean shift taking place at the time TB (Time Break) and diet is the corresponding trend variable. DU_t = 1, if $t > TB$ and Zero otherwise. Equally, DT_t = t-TB *if* $t > TB$ and Zero otherwise.

3.3 ARDL Bound cointegrtaion test

Pesaran *et al.,* (2001) developed the ARDL bound test cointegration approach. This cointegration testing technique supports in ascertaining whether the underlying series have long run relationship or not. The ARDL $(p, q_{1,} q_{2\ldots} q_k)$ model is specified as:

$$\phi(L, p) y_t = \sum_{i-1}^{k} \alpha_i (L, q_i) x_{it} + \gamma w_t + \varepsilon_t \text{ , where}$$
$$\phi(L, p) = 1 - \phi_1 L - \phi_2 L^2 - \ldots \phi_p L^p$$
$$\alpha(L, q) = 1 - \alpha_1 L - \alpha_2 L^2 - \ldots - \alpha_q L^q \quad \text{for } i = 1,2,3\ldots\ldots.k, \quad \varepsilon_t \sim iid(0; \delta^2)$$

where L indicates the lag operator such that $L^0 y_t = X_t, L^1 y_t = y_{t-1},$ and w_t is a k ×1 vector of deterministic variables with the fixed lags (see, Nkoro and Uko, 2016 for details).

3.4 Toda-Yamamoto Granger causality test

The study employ Toda and Yamamoto (1996) Granger causality test. This test has the following advantages. First, it pay no attention to any possible non-stationary or cointegration among variables when testing for causality. In other word, it does not require cointegration test, however, a likely pretest bias is avoided (Wolde-Rufael, 2005). Second, it allows the estimation of VAR at levels regardless of wheather the variables follow the same order of integration or mixed. Hence, no loss of information due to variable differencing and the method is adjustable regarding the variables at levels (Hatemi–J and Irandoust, 2002). Third, it is appropriate for finite samples since developing countries lack quality data (Nepal and Paija, 2019). However, to estimate the T-Y Granger causality, in the first place, we determine the maximum lag length (k) of a VAR in levels by lag order selection criteria. Second, testing the VAR (k+d) model as specified:

$$v_t = \alpha_v + \beta_1 v_{t-1} + \beta_2 v_{t-2} + \beta_3 v_{t-3} + \ldots + \beta_k v_{t-k} \ldots + \beta_{k+d} v_{t-k-d} + \mu_{vt} \quad (6)$$

where $V_t = (rer_t, olp_t, gex_t, cpi_t);$ αv is a (4×1) vector of constant, $\beta_{1,2,3\ldots k+d}$ are (4×4) coefficients matrix and μ_{vt} (4×1) vector indicates the stochastic term.

4. Results and Discussion

Table 1 presents the output of the unit root tests. Panel A reports the ADF unit root results where ln*rer* and ln*olp,* are not stationary at level afterward they become stationary at first difference i.e. $I(1)$ process. While ln*gex* and *ifn* are found to be stationary at level i.e. $I(0)$. Panel B presents the PP unit root test results which provide additional evidence for the mixed order of integration. Following these results, we aptly conclude that the underlying series have mixed order of integration.

Table 1: Unit root tests

Variable	Test for $I(0)$		Test for $I(1)$		I(d)
	t-statistics	Prob.	t-statistics	Prob.	
Panel A: Augmented Dickey Fuller (ADF) test					
$\ln rer_t$	-2.999	0.143	-5.237	0.000***	$I(1)$
$\ln olp_t$	-2.908	0.169	-6.311	0.000***	$I(1)$
$\ln gex_t$	-4.925	0.001***	-14.745	0.000***	$I(0)$
ifn_t	-3.986	0.015**	-7.041	0.000***	$I(0)$
Panel B: Phillips Perron (PP) test					
$\ln rer_t$	-2.479	0.336	-5.168	0.000***	$I(1)$
$\ln olp_t$	-2.924	0.164	-6.313	0.000***	$I(1)$
$\ln gex_t$	-5.328	0.000***	-15.441	0.000***	$I(0)$
ifn_t	-3.235	0.089*	-14.987	0.000***	$I(0)$

Notes: ***, ** and * indicate 1%, 5% and 10% significance level.

Table 2 presents the Zivot and Andrews unit root with structural breaks. The results suggest mixed order of integration with no bias. The results also indicate that the time break for *rer, olp*, *gex* and *ifn* falls in 1987, 1986, 2004 and 1996 respectively. The structural breaks in *rer* and *olp* coincides with the fluctuations of oil prices in the mid-1980s when OPEC attempt to stabilize the oil prices through production quotas, instead, the unfair quota policy created further increase in oil prices, which create a more than proportionate appreciation in exchange rate. This suggests that exchange rate appreciation is associated with changes in oil prices (Ogundipe *et al.,* 2014). In comparison, structural breaks in *gex* and *ifn* correspond to the era of high oil revenues leading to a rapid rise in government spending due to inflation and of real exchange rate appreciation.

Table 2: Zivot and Andrew unit root test

Variable	Test for $I(0)$		Test for $I(1)$		$I(d)$
	t-statistics	T_B	t-statistics	T_B	
$\ln rer_t$	-2.784	1985	-6.751***	1987	$I(1)$
$\ln olp_t$	-3.623	2005	-8.425***	1986	$I(1)$
$\ln gex_t$	-6.518***	2003	-6.219***	2004	$I(0)$
$\ln ifn_t$	-5.153*	1995	-7.436***	1996	$I(0)$

Note: *** and * indicate 1% and 10% significance level.

Having confirmed that the variables have mixed order of integration, we employ the Autoregressive Distributed Lag model (ARDL) bound testing approach. This approach has numerous advantages. First, it provides short-run and long-run coefficient estimates simultaneously. Second, the long-run and short run coefficients are also the basis of the cointegration test since they include $I(0)$*,* $I(1)$ or a combination of both. Third, it is appropriate and unbiased particularly when dealing with finite samples (Pesaran *et al.*, 2001). The ARDL bound testing involves two steps. First, to identify the long run relationship using the

Wald test (*F*-test). The null hypothesis is that the coefficients of the lagged variables are equal to null, which imply the absence of a long term relationship, whereas the alternative hypothesis means that at least one of the coefficients is not equal to zero. The null and alternative hypothesis in symbolic terms are:

Null hypothesis : $H_0 : \varphi_1 = \varphi_2 = \ldots \varphi_k = 0$ (No cointegration exists)
Alternative hypothesis: $H_1 : \varphi_1 \neq \varphi_2 \neq \ldots \varphi_k \neq 0.$ (Cointegration exists)

The computed *F*-statistics is compared with critical values provided by Pesaran or Narayan (Pesaran *et al.*, 2001; Narayan, 2005). If the value *F*-statistics is higher than the upper limit value, then we infer that the alternative hypothesis that there is a long term relationship is not rejected. At the other hand, if the computed *F*-statistics is less than the lower limit value, the null hypothesis is accepted which implies there is no long term relationship. Second, once cointegration relation is established, the next step is to estimate the long run and short run coefficients using lag order selection base on Akaike Information Criteria (AIC) or Schwatz Bayesian Criteria (SBC) (Adamu and Rajah, 2016). Following Pesaran *et al.,* (2001), the ARDL model corresponding to Eq. (2) is specified as:

$$\begin{aligned}\Delta \ln rer_t &= \alpha_0 + \alpha_1 \ln rer_{t-1} + \alpha_2 \ln olp_{t-1} + \alpha_3 \ln gex_{t-1} + \alpha_4 ifn_{t-1} \\ &+ \sum_{i=1}^{k} \alpha_{5,t} \Delta \ln rer_{t-1} + \sum_{i=0}^{k} \alpha_{6,t} \Delta \ln olp_{t-1} + \sum_{i=0}^{k} \alpha_{7,t} \Delta \ln gex_{t-1} + \\ &\sum_{i=0}^{k} \alpha_{8,t} \Delta ifn_{t-1} + \varepsilon_{1,t}\end{aligned}$$

Following equation (7), we set up the error correction representation model to attain the short run coefficients as follows:

$$\begin{aligned}\Delta \ln rer_t &= \alpha_0 + \sum_{i=2}^{k} \alpha_{1,t} \Delta \ln rer_{t-1} + \sum_{i=0}^{k} \alpha_{2,t} \Delta \ln olp_{t-1} + \sum_{i=0}^{k} \alpha_{3,t} \Delta \ln gex_{t-1} + \\ &\sum_{i=0}^{k} \alpha_{4,t} \Delta ifn_{t-1} + \eta ect_{t-1} + \varepsilon_{1,t}\end{aligned} \tag{8}$$

The *F*-statistics test is sensitive to the number of lags, therefore, 3 lags were chosen based on Akaike Information Criterion (AIC). Table 3 presents the results of the computed ARDL cointegration test. The computed *F*-value of 5.856 is greater than the upper critical bound value at the 1 percent significance level. This makes our conclusion valid by rejecting the null hypothesis against the alternative hypothesis that the long run cointegration relationships among the variables exists.

Table 3: ARDL Co-integration test

Model	Lag	*F*-stat	Decision	
Real exchange rate (*rer*)	3	5.856	Cointegrated	
Asymptotic critical value				
	Pesaran *et al.*, (2001)		Narayan (2005)	
Level of significance	LBV, *I*(0)	UBV, *I*(1)	LBV, *I*(0)	UBV, *I*(1)
1%	4.29	5.61	4.865	6.360
5%	3.23	4.35	3.500	4.700
10%	2.72	3.77	2.873	3.973

Notes: LBV and UBV denotes Lower and Upper Bound Values.

Table 4 reveals the long run coefficients. Except inflation, oil price and government expenditure are statistically significant at most 10 percent level. The coefficient of oil price (ln*olp*) is 0.635, positive and statistically significant at the 5 percent level. This implies that a 1 percent increase in the oil price would appreciate real exchange rate roughly by 61 percent. It is worth to mention that this result is consistent with *a priori expectation* and coincides with other studies such as Olagunju, (2005); Dulger *et al.,* (2013); Wyzan, (2005); Looney, (1991); Adenikinju *et al.,* (2002) and contradicts the finding of Sala-i-Martin (2013). This finding indicates that an increase in oil price contributes significantly to appreciation of real exchange rate. Also, it shows that Nigeria has experienced high pressure on the domestic currency as resource-abundant country, primarily oil and gas, and has responded to the growing demand for oil, which translates the prevalence "Dutch disease". At the point of 1 percent, government expenditure is positive and statistically significant at the 1 percent level. This implies that an increase of 1 percent in government expenditure would lead to an appreciation of around 91 percent in the real exchange rate. The result indicates that real exchange rate appreciates in response to an increase in government expenditure. On the other hand, the variable inflation does not present a definite sign, and statistically insignificant.

Tabel 4: Long run coefficients;
dependent variable = ln*rer*

Variable	Coefficients	Std. Error	*t*-statistics	Prob.
Constant	3.821	0.950	4.019	0.000***
$lnolp_t$	0.635	0.287	2.214	0.034**
$lngex_t$	0.905	0.239	3.784	0.000***
ifn_t	-0.015	0.011	-1.315	0.198

Note: *** and ** denotes 1% and 5% significance level.

Table 5 reports the error correction representation of the ARDL model in first difference. Suprisingly, a positive and significant lagged real exchange rate is observed implying that real exchange rate is a persistent variable in response to its determinants. The coefficients of oil price is positive and statistically significant at 1 percent level. This suggest that 1 percent increase in oil price would lead to appereciation of real exchange rate approximately by 33 percent. Likewise, a 1 percent rise in inflation would lead to appreciation of real exchange rate by 2 percent. Government expenditure is positive as expected and significant at 1 percent level. This indicates that a 1 percent rise in government expenditure will appreciate real exchange rates by 42 percent. The result also establish that appreciation of exchange rate withstand a proportionate increase in government expenditure. The coefficient of inflation is positive and significant. In principle, excess money supply puts upward pressure on prices of non-tradable goods, and leads to inflation and real exchange rate appreciation. This implies that a rise in inflation is associated with appreciation in real exchange rate, and studies have reached the same conclusion (see, Jbir and Zouari-Ghorbel, 2011; Kutan and Wyzan, 2005). The ect_{t-1} is the error correction term which measures the speed of adjustment towards the long run equilibrium, which found to be -0.463 and significant as expected, which implies that there is cointegration among the varibles. Moreover, a short run deviation to the long run equilibrium is corrected by 46 percent in the event of any shock in the system. Overall, the entire findings show that all variables of oil prices, government expenditure and inflation plays a major role in determining real exchange rate appreciation in Nigeria.

Nevertheless, obtaining the estimated short run model alone could not ensure the reliability and stability of the model. Bearing this in mind, we subjected the dynamic model into an appropriate diagnostic tests to ensure that the model do not violate the econometric assumptions. Table 6 reports the results of the diagnostic tests. The langragian multiplier test for serial correlation shows the absence of autocorrelation in stochastic error terms, the Jaque-bera test for normality of the residuals also proves that the errors are normally distributed. Furthermore, there is no evidence of autoregressive conditional heteroscedisticity (ARCH). The Ramsey RESET for functional form also confirm that the model is correctly specified. Finally, Pesaran and Shin (1999) suggests the use of cumulative sum (CUSUM) and the cumulative sum of squares (CUSUMSQ) of the recursive residual test proposed by Brown, Durbin and Evans (1975) in testing the model structural stability. The cusum and cusum square (see Figure 2) indicates that the test statistics are within the critical band at the 5% confidence interval, thus, suggesting that the overall coefficients over the sample period are generally stable. The *R*-squared and Adjusted *R*-squared of 90 percent and 85 percent are reasonable and well behaved to explain the variables relation.

Tabel 5: Error correction representation coefficients
dependent variable = ln*rer*

Variable	Coefficients	Std. Error	*t*-statistics	Prob.
$\Delta \ln rer_{t-1}$	0.276	0.147	1.876	0.070*
$\Delta \ln rer_{t-2}$	0.441	0.155	2.837	0.008***
$\Delta \ln rer_{t-3}$	0.374	0.156	2.386	0.023**
$\Delta \ln olp_{t}$	0.046	0.152	0.303	0.763
$\Delta \ln olp_{t-1}$	0.327	0.109	2.989	0.004***
$\Delta \ln gex_{t}$	0.419	0.135	3.093	0.004***
Δifn_{t}	0.003	0.003	1.013	0.319
Δifn_{t-1}	-0.001	0.004	-0.237	0.813
Δifn_{t-2}	0.015	0.003	4.178	0.000***
Speed of adjusment				
ect_{t-1}	-0.463	0.136	-3.400	0.002***
R^2	0.900			
Adj. R^2	0.852			
F-stat (Prob.)	18.748 ***			
DW	1.94			

Note: ***, ** and * denotes 1%, 5% and 10% significance level.

Table 6: Diagnostic check

Test	*F*-value	Prob.
Serial correlation $\chi^2(2)$	0.002	0.997
Normality test χ^2	0.511	0.774
ARCH $\chi^2(1)$	0.238	0.627
Ramsey RESET χ^2 (1)	0.312	0.579

Nevertheless, ARDL cointegration approach indicates whether or not the long run cointegration exists among the variables but it does not reveal the direction of causality. For this reason, we use the modified Wald test (MWALD) suggested by Toda and Yamamoto (1995) to determine the long run causal relation among *rer*, *olp*, *gex* and *ifn.* Now, to investigate the T-Y multivariate causality, the following system of equations for VAR are specified:

$$\ln rer_t = \alpha_0 + \sum_{i=1}^{k} \alpha_{1i} \ln rer_{t-i} + \sum_{j=k+1}^{d\max} \alpha_{2j} \ln rer_{t-j} + \sum_{i=1}^{k} \delta_{1i} \ln olp_{t-i} + \sum_{j=k+1}^{d\max} \delta_{2j} \ln olp_{t-j}$$

$$+ \sum_{i=1}^{k} \phi_{1i} \ln gex_{t-i} + \sum_{j=k+1}^{d\max} \phi_{2j} \ln gex_{t-j} + \sum_{i=1}^{k} \vartheta_{1i} ifn_{t-i} + \sum_{j=k+1}^{d\max} \vartheta_{2j} ifn_{t-j} + \varepsilon_{1,t} \quad (9)$$

$$\ln olp_t = \beta_0 + \sum_{i=1}^{k}\beta_{1i}\ln olp_{t-i} + \sum_{j=k+1}^{d\max}\beta_{2j}\ln olp_{t-j} + \sum_{i=1}^{k}\varphi_{1i}\ln rer_{t-i} + \sum_{j=k+1}^{d\max}\varphi_{2j}\ln rer_{t-j}$$
$$+ \sum_{i=1}^{k}\gamma_{1i}\ln gex_{t-i} + \sum_{j=k+1}^{d\max}\gamma_{2j}\ln gex_{t-j} + \sum_{i=1}^{k}\varsigma_{1i} ifn_{t-i} + \sum_{j=k+1}^{d\max}\varsigma_{2j} ifn_{t-j} + \varepsilon_{2,t} \qquad (10)$$

$$\ln gex_t = \theta_0 + \sum_{i=1}^{k}\theta_{1i}\ln gex_{t-i} + \sum_{j=k+1}^{d\max}\theta_{2j}\ln gex_{t-j} + \sum_{i=1}^{k}\vartheta_{1i}\ln rer_{t-i} + \sum_{j=k+1}^{d\max}\vartheta_{2j}\ln rer_{t-j}$$
$$+ \sum_{i=1}^{k}\sigma_{1i}\ln olp_{t-i} + \sum_{j=k+1}^{d\max}\sigma_{2j}\ln olp_{t-j} + \sum_{i=1}^{k}\eta_{1i} ifn_{t-i} + \sum_{j=k+1}^{d\max}\eta_{2j} ifn_{t-j} + \varepsilon_{3,t} \qquad (11)$$

$$ifn_t = \theta_0 + \sum_{i=1}^{k}\theta_{1i} ifn_{t-i} + \sum_{j=k+1}^{d\max}\theta_{2j}\ln \inf_{t-j} + \sum_{i=1}^{k}\vartheta_{1i}\ln rer_{t-i} + \sum_{j=k+1}^{d\max}\vartheta_{2j}\ln rer_{t-j}$$
$$+ \sum_{i=1}^{k}\sigma_{1i}\ln olp_{t-i} + \sum_{j=k+1}^{d\max}\sigma_{2j}\ln olp_{t-j} + \sum_{i=1}^{k}\nu_{1i}\ln gex_{t-i} + \sum_{j=k+1}^{d\max}\nu_{2j}\ln gex_{t-j} + \varepsilon_{4,t} \qquad (12)$$

Table 7 reports the T-Y multivariate Granger causality tests of equation 9, 10, 11 and 12. The independent variables in each single equation of VAR have their *F*-statistics, which support the short run causality for the respective pair of variables. It tends to be seen that the result indicate a unidirection causality running from oil price to real exchange rate. Like the dynamic model, causality test firmly established that higher oil price is foundamental factor for real exchange rate appreciation. Besides, bidirectional Granger causality running between real exchange rate and inflation, oil price and government expenditure. It should be noted that there is no evidence of unidirectional or bidirectional causality running from inflation to government expenditure and oil prices, and oil prices to inflation, real exchange rates and government expenditure. Hence, improving significant capital expenditure to lagging sector particularly agriculture will be a priority on minimizing the effect of Dutch disease in Nigeria.

Table 7: Toda-Yamamoto Multivariate Granger causality test

Variable	*rer*	*olp*	*gex*	*ifn*	*Decision*
rer	-	3.437 (0.070)*	0.886 (0.346)	2.843 (0.091)*	*olp* → *rer*; *ifn* → *rer*
olp	0.746 (0.387)	-	7.750 (0.005)***	0.009 (0.922)	*gex* → *olp*
gex	2.410 (0.120)	5.213 (0.022)**	-	1.532 (0.215)	*olp* → *gex*
ifn	3.192 (0.074)*	0.379 (0.537)	0.001 (0.971)	-	*rer* → *ifn*

Notes: ***, ** and* indicates 1, 5 and 10 percent level of significance. T-Y causality tests are estimated.

5. Conclusion and Policy implications

This paper explores the empirical analysis for validating the Dutch disease in oil-rich Nigeria. Using the autoregressive distributed lag approach, the empirical results reveal a number of interesting findings. First, the Dutch disease model indicates a cointegrating relationship among the variables. Second, there is clear evidence that oil price, government expenditure and inflation have significant impact on real effective exchange rate. This confirms the existence of Dutch disease in Nigeria. Third, the result from the causality test shows a unidirectional causality running from oil price to real exchange rate, and bidirectional causality between oil price and government expenditure, on one hand real exchange rate and inflation.

However, this study draws three important lessons from the policy viewpoint. First, government could increase investment in the non-tradable sectors of the economy to minimize heavy dependence on oil revenues that are capable of triggering a real exchange rate appreciation. Second, along with the first recommendation, it is particularly important for government to focus on investment in human resources, basic infrastructure projects and tax concessions to raise the level of competitiveness of the lagging sectors of the economy. Third, we may expect government to shift to inflation targeting policy in order to lower loan prices by minimzing inflation, thus, enhancing their accessibility to nontradeable sectors. Finally, the framework adopted by this paper is idiosyncratic and should not be regarded as general, however, there are variant factors in explaining Dutch disease in Nigeria and across the minerals-based countries. Therefore, in order to consider other likely factors, which are missing from this paper, further analysis must be carried out.

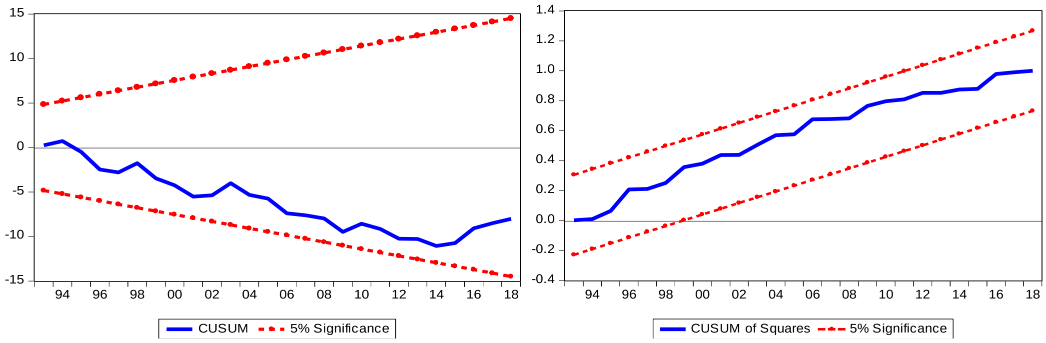

Figure 2: Residual plots for CUSUM and CUSUMSQ stability test

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