AFRICAN JOURNAL OF ECONOMIC POLICY

Vol. 20, No. 2, Dec. 2013
ISSN 1116-4875

Vol. 20, No. 2, Dec. 2013

AN EMPIRICAL RE-EXAMINATION OF EXCHANGE RATE-TRADE BALANCE NEXUS IN NIGERIA *

Mutiu A. Oyinlola
Department of Economics, University of Ibadan, Nigeria

Olusegun A. Omisakin
Department of Economics and Business Studies, Redeemer's University, Nigeria; and Centre for Econometric and Allied Research (CEAR), University of Ibadan, Nigeria

Oluwatosin A. Adeniyi
Department of Economics, University of Ibadan, Nigeria

ABSTRACT
The Nigerian exchange rate-trade balance nexus was re-examined. The long run relationship between these variables was explored using the Gregory-Hansen cointegration approach on a data sample between 1980:Q1 and 2010:Q4. Prior to this, three efficient integration tests that can overcome potentially severe finite sample power and size problems suffered by the standard methods were tactfully pursued for robustness. The short run impact analysis was done in the error correction framework. The analyses showed that exchange rate depreciation led to trade balance deterioration in both the short run and the long run. Thus, this study could not find support for J-curve in Nigeria. Some suggestions on the way forward were put forth.
Keywords: Trade balance; Exchange rate; Autoregressive distributed lag
JEL classification: F13; F31; C32

INTRODUCTION
Exchange rate-trade flows nexus has received generous enquiries over the past decades for its significant implication in the management of the economy. Specifically, interest of policy analysts is on the dynamics of trade balance over time in periods of manoeuvred depreciation of the exchange rate. This behaviour, in theory, is such that trade balance deteriorates in the short run and subsequently improves to a level higher than the one prior to the

* Authors acknowledge the comments of anonymous referees; meanwhile we take responsibilities for any error therein.

depreciation, thus resembling a J curve. Investigations into this J curve phenomenon have traversed time and space. Vintage enquiries led by Robinson (1947) and Meltzer (1948), focused on the Marshal-Lerner condition to unravel the nature of the relationship.[1] This effort led to the idea of the development of the elasticity approach to trade balance.

A number of studies found evidence in support of J-curve while others depict non-existence and rather found other distinct effects from the J-curve. Studies such as Bahmani-Oskooee (1985), Krugman and Baldwin (1987), Carter and Pick (1989), Koray and McMillin (1999), Leonard and Stockman (2001), Singh (2002), Akbostanci (2002), Hacker and Hatemi (2004), Narayan (2004), Hsing (2005) as well as Davinda and Swaha (2013), find positive evidence in support of the J-curve effect. Bahmani-Oskooee (1985) investigated the J-curve phenomenon for Greece, India, Korea and Thailand and find evidence in favour of the J-curve effect for all the countries except for Thailand. Krugman and Baldwin (1987) established empirical evidence of a J-curve on a US data. Koray and McMillin (1999) find evidence that depicts the converse of J curve. They posit that appreciation of exchange rate initially improves the trade balance and later deteriorates it. This lends support to an inverted J-curve.

The existence or otherwise of J-curve is predicated on the emergence of long run relationship between exchange rate and trade balance incidental on the stationarity of the series in question. Substantial part of the studies on J-curve depends on the first generation approaches of unit root testing; viz: Augmented Dickey-Fuller (ADF) (1979) and Phillips and Perron, (PP) (1988).

---

[1] The ML condition holds that devaluation of a country's real exchange rate should improve her trade balance, if the sum of the country's price elasticity of demand for exports and imports are greater than one in absolute value. Devaluation of the exchange rate causes the price of exports to become cheaper, the demand for exports will therefore increase. However, the price of imports will rise, and the demand for imports will decrease. The net effect on the trade balance will depend on price elasticities. If exports are elastic to price, their demand will increase proportionately more than the decrease in price, and the total export revenue will increase. If imported goods are elastic, the total import expenditure will decrease. The trade balance will improve in both cases.

Glynn *et al.* (2007) submit that methods of estimation of economic relationships and modelling fluctuations in economic activities have been subjected to fundamental changes. The method of estimation of the standard regression model, Ordinary Least Square (OLS), is based on the assumption that the means and variances of these variables being tested are constant over time. Variables whose means and variances change over time are known as non-stationary or unit root variables. Therefore, incorporating non-stationary or unit root variables in estimating the regression equations using OLS method give misleading inferences. Instead, if variables are non-stationary, the estimation of long run relationship between those variables should be based on the cointegration method. Since the testing of the unit roots of a series is a precondition to the existence of cointegration relationship, originally, the Augmented Dickey-Fuller (ADF) (1979) test was widely used to test for stationarity. Notably, Perron (1989) shows that the failure to allow for an existing break leads to a bias that reduces the ability to reject a false unit root null hypothesis. The powers of these tests (ADF and PP) are known to suffer potentially severe finite sample power and size problems (see for instance, DeJong, *et al.*, 1992; and Schwert, 1989). To overcome this, Perron (*op. cit.)* proposes a model allowing for a known or exogenous structural break in the Augmented Dickey-Fuller (ADF) tests. Following this development, many authors, such as Zivot and Andrews (1992) and Perron (1997) suggest determining the break point *endogenously* from the data.

Devaluation engenders a long run improvement in the trade balance, provided the ML conditions are met. Given that the short run elasticities are usually smaller than the long run elasticities, the trade balance may not improve in the short run. As Magee (1973) points out, the J-Curve phenomenon is theoretically ambiguous and empirical evidence has been rather mixed or inconclusive. Various reasons ranging from the use of different techniques to different model specifications have been ascribed to this lack of conclusive result. Most studies on J-curve have focused mainly on developed economies as compiled by Bahmani-Oskooee and Ratha (2004). Nasir *et. al.* (2009) highlight three facts from the literature:

First, existing literature provides mostly inconclusive evidence on the issue of response of trade balance to exchange rate shock. Second, the studies investigating developing countries' samples are limited in number and coverage. Third, studies such as Bahmani-Oskooee and Brooks, 1999; Bahmani-Oskooee and Ratha, 2004; Bahmani-Oskooee and Tatchawan, 2001; utilising a recent development, autoregressive distributed lag approach, in the cointegration literature, appears to identify some type of adjustment in the trade balance following currency depreciation. The present effort therefore hopes to increase the spread of the coverage of the investigation by investigating the existence of the phenomenon in Nigeria with the use of second generation stationarity tests and the Zivot and Andrew (1992) structural break unit root test. The long run relationship among variables is unravelled by the Gregory-Hansen structural break cointegration.

Objective of the Study

The argument in the literature is that a real depreciation initially deteriorates the trade balance but over time, the trade balance improves such that the time path associated with the response of the trade balance generates a tilted J-curve. Establishing the relationship between trade balance and exchange rate for Nigeria is therefore the main objective of this study. Specifically, the study seeks to determine the long run and the short-run impacts of real exchange rate depreciation on trade balance in Nigeria between 1980:Q1-2010:Q4 using advances in time-series econometrics.

Justification of the Study

This study adopts the model of Rose and Yellen (1989) used by many other empirical studies in the literature (Wilson, 2001; Baharumshah, 2001; Onafowora, 2003). The model suggests that bilateral trade balance is influenced by real exchange rate, domestic and foreign incomes. The present effort is distinct in two main ways: first, the treatment of the data series is detailed and unparalleled. Previous studies on Nigeria known to the authors, have not considered structural breaks in the data employed. We depart from this by employing the Zivot and Andrews (ZA) (1992) tests to search for endogenous breaks. Given that they argue that the break points should be viewed as being correlated with the data, selecting the break exogenously could lead to an over rejection of the unit root hypothesis. It is also well-known from the literature on structural breaks; tests for cointegration in the presence of a break tend to under-reject the null of no cointegration (Gregory, Nason and Watt, 1996). Added to the first point is the second which dwells on the use of two other long run estimators, the fully modified OLS (FMOLS) and dynamic OLS (DOLS), known to provide robust results in small samples sizes, aside the ordinary least squares method.

The rest of the study is organised as follows. Section 2 presents a brief background, followed by literature review in Section 3. Section 4 dwells on the theoretical framework and methodology in this study. The empirical results are discussed in section 5. Finally, section 6 provides the summary and conclusion.

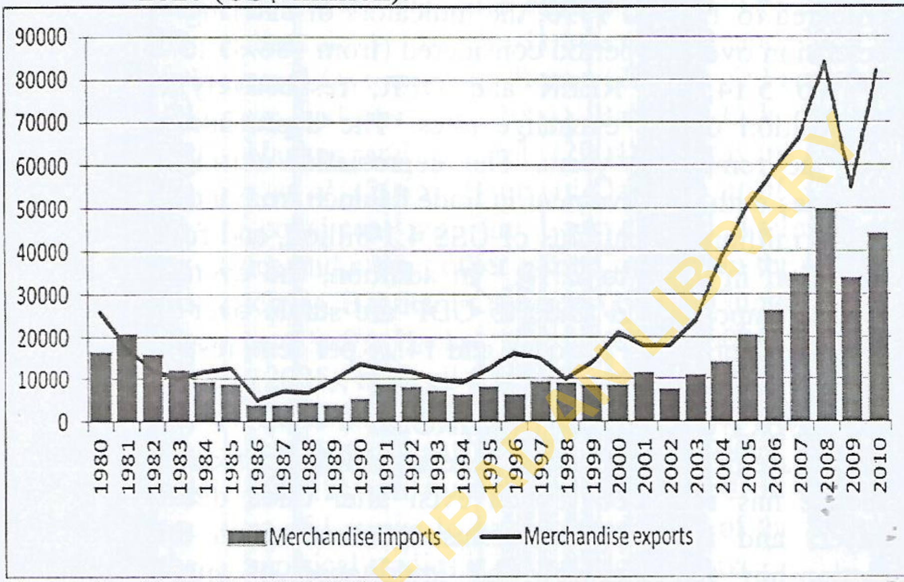
STYLISED FACTS ON TRADE FLOWS AND EXCHANGE RATE

Merchandise trade flows in Nigeria often reflect the sequence of various policies employed over time. The problems of balance of payments in the 1960s prompted the application of trade policies in the 1970s. Consequently, as expected, volume of imports decreased due to increase in import prices, being the outcome of high tariff rates on imported goods. Coupled with this was the use of administratively determined exchange rate to ensure cheap import of raw materials for local manufacturing industries; nonetheless; the problems persisted.

The unabated problems of balance of payments, the fall in price of crude oil and the attendant economic depression, heralded the introduction of Structural Adjustment Programme (SAP) in 1986. As a result, trade and exchange rate policies were liberalised. Import restrictions were reduced and export prohibition abolished. Although, the tariff structure was modified 1995, it remained liberalised.

Figure 1 shows that merchandise imports and exports fell considerably between 1980 and 1985. This was the period of economic recession and chronic balance of payments problems precipitated by the oil price crash. The perturbing spots can easily be detected from the Figure in the year 1981, 1982 and 1983, whereby merchandise imports surpassed merchandise exports, thus it led to trade deficit (Table 1). Thereafter, SAP was introduced in 1986 to address the problem, among others. The impact of this intervention is evident as exports witnessed marginal increase between 1988 and 1990. Trade performance subsequently, was not impressive due to encumbrances ranging from political instability to policy reversals. The trend however changed at the inception of the democratic governance in 1999. The perception of political stability engendered by democracy as well as the vigorously pursued liberalisation policy at that time could be responsible. The five-year average value of exports rose gradually from USD 8.6 billion (1986-1990) to USD 29.8 billion (2001-2005). The country's exports rose by 23.5 per cent in 2008 to an estimated US$ 76.3 billion from US$ 61.8 billion in 2007; the average value in 2006 to 2010 is about US$ 70 billion. This has been attributed to higher prices of oil exports during the first half of 2008. Imports on the other hand, averaged US$ 4.5 billion in 1986 to 1990, reaching about US$ 13 billion in 2001to 2005. Indeed, both imports and exports rose drastically until 2009 when they fell due to the global financial crisis. The trend imports had since picked up appreciably since the crisis waned. For example, exports was put at US$ 7.93 billion in September of 2012 from US$ 8.67 billion in August of 2012; while imports increased to US$ 4.55 billion in September of 2012 from US$ 3.98 billion in August of 2012.

Figure 1: Merchandise Imports and Exports for Nigeria, 1980-2010 (US$'million).



Source: World Bank's WDI, 2012.

Table 1: Five-year Average Trend in Exchange Rates and Trade Flows in Nigeria, 1980 to 2010.

Year	Imports	Exports	Trade balance	REER	OER	Import/ GDP	Export /GDP	Trade/ GDP
1980	16660.00	25968.00	9308.00	343.40	0.55	25.95	40.45	14.50
1981-1985	13486.60	12958.20	-528.40	488.85	0.74	33.33	34.03	0.70
1986-1990	4495.40	8598.20	4102.80	136.85	5.14	18.92	35.34	16.42
1991-1995	7920.80	11163.00	3242.20	82.13	18.63	30.11	42.27	12.16
1996-2000	8491.80	15209.20	6717.40	117.59	51.94	23.36	40.77	17.41
2001-2005	12980.80	29829.80	16849.00	90.73	125.04	17.51	38.49	20.98
2006-2010	37889.01	69289.99	31400.98	111.01	134.44	21.22	39.14	17.92

Source: Computed from the World Bank's WDI, 2012.

Notes: REER- Real Effective Exchange Rate, OER-Official Exchange Rate (naira per US$); values of imports, exports and trade balance in US$ million, while ratios are in percentages.

When the period prior to liberalisation, 1981to 1985, is compared to 1986 to 1990, the indicators of exchange rates show declension over the period considered (from 488.85 to 136.85, and 0.74 to 5.14, for REER and OER, respectively), indicating depreciation of the exchange rates. The depreciation continues over the remaining years. The depreciation during the period coincides with improvement in trade balance from a deficit of US$ 528.40 million to a surplus of US$ 4.1 billion, and further to US$ 31 billion in 2006 to 2010. In addition, the average ratios of exports, imports, and trade to GDP are stable over the years at 38.34 per cent, 24.04 per cent, and 14.26 per cent, respectively.

SUMMARY OF THE LITERARURE

The debate of the influence of exchange rate depreciation on trade balance has remained topical even after three decades. Policy makers and indeed governments have relied on this essential relationship to argue on the importance of macroeconomic policies, geared towards achieving the goals of the economy. Theoretically, exchange rate depreciation affects the trade balance through two main channels. The volume effect, due implicitly to more expensive imports, raises the export volume while the volume of import declines resulting in an increase in the trade balance. A countervailing import value effect moves the trade balance in the opposite direction (Krugman and Obstfeld, 2001; Hacker and Hatemi, 2004). Hence, the net effect on the trade balance depends on which effect dominates. There are expectations that deficit in the trade balance of a country may be eliminated by a real devaluation in the domestic currency. Success of devaluation, however, depends on whether or not the sum of import and export elasticities exceeds unity, also known as the Marshall- Lerner (ML) condition. Bahmani-Oskooee (1985) contends that there have been cases under which the ML condition was satisfied yet the trade balance continued to deteriorate. Therefore, he concludes that the focus of a trade policy should be on the short run dynamics that trace the post devaluation time path of the trade balance implying the J-curve phenomenon.

There are quite a number of studies that have investigated J-curve phenomenon over the past decades. A comprehensive survey of 37 studies on the J-curve between 1973 and 2003 is provided by Bahmani-Oskooee and Ratha (2004a). Other recent studies are Narayan and Narayan (2004), Bahmani-Oskooee and Ratha (2004b), Hacker and Hatemi (2004), Narayan (2004), Bahmani-Oskooee *et al.* (2005), Bahmani-Oskooee *et al.* (2006). Some results are consistent with the J-curve trend while others depict non-existence and reflect other effects other than the J-curve effect. Studies such as Bahmani-Oskooee (1985), Gupta-Kapoor and Ramakrishnan (1999), Hacker and Hatemi (2004), Narayan (2004) and Hsing (2005), have all find positive evidence in support of the J-curve effect. Others such as Koch and Rosensweig (1990), Bahmani-Oskooee and Goswani (2003), Khatoon and Rahman (2009), Simakova (2012) and Awan *et al.* (2012) find little or no evidence in support of J-curve. In sum, the evidence of the J-curve offered by these empirical studies is at best inconclusive.

Only a sparse number of studies on J-curve in Nigeria exist. Bahmani-Oskooee and Gelan (2012) who included Nigeria in the test of the phenomenon for nine African countries, could not establish support for J-curve. Oyinlola *et al.* (2013) using three cointegration techniques (Engle-Granger, Johansen, and Autoregressive distributed lag) likewise find no support for J-curve. Enquiries into J-curve explore various dimensions to the argument; ranging from the observation of the phenomenon at a more disaggregated level, use of intervening variables, measurement issues, as well as methods. The present effort is an attempt to explore, the existence of J-curve in Nigeria with the use of efficient methods of testing integration of series and long run relationship that account for structural break in the series. This is a pioneer study, to the best of our knowledge, in this direction on Nigeria.

THEORETICAL FRAMEWORK AND METHODOLOGY

Analytical Framework

Rose and Yellen (1989) framework began with a specification of the import demand equations. As in Marshallian demand analysis, the volume of imported goods demanded by the home (foreign) country is determined by real domestic (foreign) income and the relative price of imported goods. Clearly, real income has a positive impact on the volume of import demand and the relative price of imported goods has a negative relationship. Demand for imports is given as:

$$D_m = D_m(YN, p_m) \text{ and } D_m^* = D_m^*(YW, p_m^*) \qquad (1)$$

Where: $D_m(D_m^*)$ is the quantity of goods imported by the home (foreign) country; $YN(YW)$ is the level of real income measured in domestic (foreign) output; p_m is the relative price of imported goods to domestically produced goods, both measured in home currency; and p_m^* is the analogous relative price of imports abroad.

Likewise, they specify the equations for the supply of exportables. In a simple purely competitive market, the relative price of exportables determines their supply, expressed as:

$$S_x = S_x(p_x) \text{ and } S_x = S_x^*(p_x^*) \qquad (2)$$

Where: $S_x = (S_x^*)$ is the supply of home (foreign) exportables; p_x is the home country relative price of exportables (defined as the ratio of the domestic currency price of exportables p_x to the domestic price level, P); p_x^* is defined as the foreign currency price of exportables, p_x^* divided by the foreign price level, P^*.

Therefore, the domestic relative price of imports can be written as:

$$p_m = EP_x^* / P = REERP_x^* \qquad (3)$$

Where: E is the nominal exchange rate, defined as the number of domestic currency units per unit of foreign currency, and REER is the real effective exchange rate. Similarly, the relative price of imports abroad is:

$$p_m^* = p_x / REER \tag{4}$$

In equilibrium, quantities of trade and the relative price of exported goods in each country are determined by the two equilibrium conditions:

$$D_m = S_x^* \text{ and } D_m^* = S_x \tag{5}$$

The value of the home country's balance of trade in real terms is the difference between the value of exports and the value of imports in domestic currency:

$$B = p_x D_m^* - REER.p_x^*.D_m \tag{6}$$

Equations 1, 2,3,4,5 to 6 yield the following reduced form:

$$B = B(REER, YN, YW) \tag{7}$$

Rose and Yellen (1989), and Aziz (2008) estimated a log-linear variant of Equation 7:

$$TB_t = a + b\ln YN_t + c\ln YW_t + d\ln REER_t + \varepsilon_t \tag{8}$$

Where: TB_t is the Nigeria trade balance with the rest of the world, usually measured as the difference between the value of total exports and imports. In this study we measure trade balance as the ratio of the exports value to the imports value. The X/M ratio or its inverse has been used in many empirical investigations of the trade balance-exchange rate relationship (Onafowora, 2003; and Bahmani-Oskooee and Brooks, 1999). Such a measure is not only unit free but also reflects movements of the trade balance in real and nominal terms. YN_t is the index of Nigeria real GDP, YW_t is the rest of the world's GDP, and $REER_t$ is the real effective exchange rate between Nigeria naira and the rest of the world, defined such that an increase reflects a real depreciation of the naira against the major trading currency.

It is expected that the volume of imports (exports) in a home country (foreign country) will increase with an increase in real income and purchasing power in the home country (trading partners), and vice versa. Hence, we expect $b<0$ and $c>0$. If the increase in real income is as a result of an increase in production of import substitutes, imports may fall as income rises. This implies in this case that $b>0$ and $c<0$. The impact of exchange rate changes on trade balance is ambiguous, that is, d could be positive or negative. If there is a real depreciation or devaluation of the domestic currency, that is, REER decreases, then this makes imports more expensive without a corresponding rise in export prices; the increased competitiveness in prices for the domestic country should result in it exporting more and importing less (the "volume effect"). However, the lower REER also increases the value of each unit of import (the "import value effect"), which would tend to diminish the trade balance. Krugman and Obstfeld (2001) argue that in the short run, import value effects prevail, whereas the volume effects dominate in the long run. For real exchange rate depreciation to improve trade balance, the volume effect should overwhelm the price effect. Hence, $d > 0$ satisfies the Marshall-Lerner condition[2].

[2] Marshall-Lerner condition states that devaluation will improve the trade balance only if the sum of the foreign elasticity of demand for exports and the home country elasticity of demand of imports is greater than unity. If the sum of these two elasticities is less than unity then devaluation will lead to a deterioration of the trade balance.

Methodology[3]

Unit roots

The study employed quarterly data on Nigeria from 1980:Q1 to 2010:Q4. The time series characteristics of the data is first investigated to test whether the variables are stationary or not. By definition, a time series is stationary if its means, variances and covariance are all invariant with respect to time. The most commonly used univariate unit root tests in the literature are the augmented Dickey-Fuller (ADF) (Said and Dickey, 1984) and Phillips- Perron (PP) (Phillips and Perron, 1988) tests. There is a popular view that these univariates' unit root tests have infamously low power against local stationary alternatives and suffer from serious size distortion when the data generating process (DGP) has negative moving average (MA) terms (Schwert, 1989). The efficient unit root tests developed by Elliott-Rothenberg-Stock (1996), Ng and Perron (2001) and the Zivot and Andrews (1992) were employed.[4] Elliott-Rothenberg-Stock (1996; ERS mostly henceforth) develop a feasible point optimal test that relies on local GLS detrending. This test has much greater power than standard ADF and PP unit root tests[5]. Ng and Perron (2001) extend the work done by Elliot, *et al.* (1996); and developed unit root tests based upon the local GLS detrending method, complemented with an autoregressive spectral density estimator of the long run variance. This class of tests, which they denote as the M-tests, has much less size distortion in the presence of MA errors than the standard tests (Said and Dickey, 1984).

In addition, a stationary time series may look like non-stationary when there are structural breaks in the intercept or trend.

---

[3] This section draws immensely from Narayan (2004).

[4] The presence of unit root in a variable implies that the variable is non-stationary i.e. it is integrated of order one I(1) and it has to be differenced to be made stationary, i.e. integrated of order zero I(0).

[5] Ng and Perron (2001) developed M-tests strike out as having the best size adjusted power properties. As described later Ng and Perron is an extension of the DF-GLS test of ERS (1996) and the modified Z tests of Perron and Ng (1996).

Therefore, the unit root tests could lead to false non-rejection of the null when we do not consider the structural breaks. To address this, the Zivot and Andrews (1992; ZA mostly henceforth) is employed. The central idea behind the use of ZA unit root test, is to ascertain whether, in the presence of structural breaks in the data, the series are integrated of order one or otherwise. Two versions of the ZA (1992) sequential trend break model are used to investigate the unit root hypothesis for the variables. Model A allows for a change in intercept, while model C allows for a change in the intercept and slope.[6] As an illustration, Model A has the following form:

$$\Delta y_t - \kappa + \alpha y_{t-1} + \beta t + \theta_1 DU_t + \sum_{j=1}^{k} d_j \Delta y_{t-j} + \varepsilon_t \qquad (9)$$

Model C takes the following form:

$$\Delta y_t - \kappa + \alpha y_{t-1} + \beta t + \theta_1 DU_t + \gamma_1 DT_t + \sum_{j=1}^{k} d_j \Delta y_{t-j} + \varepsilon_t \quad (10)$$

Here, Δ is the first difference operator; ε_t is a white noise disturbance term with variance σ^2; and t-1… T is an index of time. The Δy_{t-j} terms on the right-hand-side of Equations 9 and 10 allow for serial correlation and ensure that the disturbance term is white noise. Finally, DU_t is an indicator dummy variable for a mean shift occurring at time TB and DT_t is the corresponding trend shift variable, where: $DU_t - 1$ and $DT_t - t - TB$ if $t > TB$; 0 otherwise. As is conventional, *trimming region* [0.15, 0.85] is chosen and the break point where the value of TB for which the ADF t-statistic is minimised selected. The *t-sig* approach suggested by Hall (1994) is used to select the optimal lag length. The null hypothesis here is that the series [y_t] is an integrated process without a structural break, against the alternative hypothesis that [y_t] is trend

[6] Similar model variants are also specified for ERS (1996) and Ng and Perron (2001).

stationary with a structural break in the trend function which occurs at an unknown time.

Cointegration

One of the widely used methods for cointegration is the Gregory and Hansen (1996) threshold cointegration test. And the test equations are expressed as follows:

Level Shift Model: $y_{1t} = \mu_1 + \mu_2\varphi_{t\tau} + \alpha^T y_{2t} + e_t$ (11)

Level Shift and Trend Model: $y_{1t} = \mu_1 + \mu_2\varphi_{t\tau} + \beta t + \alpha^T y_{2t} + e_t$ (12)

Regime Shift Model: $y_{1t} = \mu_1 + \mu_2\varphi_{t\tau} + \alpha^T y_{2t} + \alpha^T y_{2t}\varphi_{t\tau} + e_t$ (13)

Where: y is the observed data, while μ_1 and μ_2 represent the intercept before the shift and the change in the intercept at the time of the shift; φ is the dummy variable that captures structural change;[7] β is the trend slope before the shift; α is the slope coefficients and are assumed to be constant. Y_{1t} represents the regressand while Y_{2t} denotes the independent variable(s). The standard methods of testing the null hypothesis of no cointegration are residual-based and are obtained when equations (11, 12 and 13) are estimated using OLS and the unit root tests are applied to the regression errors (Gregory and Hansen, 1996). Three tests are suggested by Gregory and Hansen (*op. cit*) in testing for cointegration with structural change (i.e. the stationarity of e_t in equations 11 to 13) among series. These statistics are the commonly used ADF statistic and extensions of the Z_α and Z_t test statistics of Phillips (1987). These statistics are defined as:

---

[7] $\varphi_t = \begin{bmatrix} 0 \ if \ t \leq (n\tau) \\ 1 \ if \ t > (n\tau) \end{bmatrix}$ where the unknown parameter $\tau \in (0,1)$ implies the timing of the break point, and $(n\tau)$ denotes integer part

$$ADF^{*} - \inf_{\tau \in T} ADF(\tau) \tag{14}$$

$$Z_{\alpha}^{*} - \inf_{\tau \in T} Z_{\alpha}(\tau) \tag{15}$$

$$Z_{t}^{*} - \inf_{\tau \in T} Z_{t}(\tau) \tag{16}$$

Equations (14), (15), and (16) are estimated sequentially with the break point changing. Non-stationarity of the obtained residuals is checked by the ADF test. Setting the test statistics to the smallest value of the ADF statistics in the sequence, we selected the value that constitutes the strongest evidence against the null hypothesis of no cointegration. In other words, the smallest values of $ADF(\tau)$, $Z_{\alpha}(\tau)$,and $Z_{t}(\tau)$ across all possible breakpoints are required to reject the null hypothesis.[8]

EMPIRICAL RESULT AND DISCUSSION

Unit root tests

The variables for our analysis are subjected to unit roots test to determine whether they are stationary or non-stationary series. Tables 2 and 3 present the tests for integration without structural break proposed by ERS and Ng-Perron while Tables 4 and 5 present the ZA test for unit roots with one structural break. All the tables contain results of model specification at the constant (intercept), and intercept (constant) and linear trend. The ERS and Ng-Perron tests suggest that Real GDP (lnYN) and exchange rate (lnREER) are each integrated of order one or $I(1)$ while trade balance (LNTB) and income of the rest of the world (LNYW) are stationary at level; though LNYW was found to be stationary at the ten per cent level with ERS. It should be noted that the order of integration of lnTB and lnYW could not be determined when Ng-Perron test was carried out, even at second difference. The issue

---

[8] see for example, Fountas and Wu, 1998 and Narayan, 2005 for details on the application

was however put to rest when the ZA test was done; the null hypothesis of non-stationarity cannot be rejected for all the series in Model C, which allows for a change in the intercept and slope at the five per cent level. This implies the series cannot be characterised as a stationary process with a break.[9]

Table 2: Elliott-Rothenberg-Stock (ERS) Test Statistic Test for Unit Root

Elliott-Rothenberg-Stock (ERS) test statistic
Null Hypothesis: LNTB has a unit root

	Constant		Constant, Linear Trend		
	level	Differenced	level	Differenced	Decision
lnTB	2.836†	0.125*	2.716*	0.347*	I(0)
lnYN	457.146	1.918*	12.611	4.589†	I(1)
lnYW	367.219	4.859	6.461‡	14.648	I(0)
lnREER	10.442	0.680*	11.529	2.352*	I(1)
Test critical values[a]:					
1% level	1.940	1.941	4.210	4.212	
5% level	3.124	3.124	5.645	5.645	
10% level	4.208	4.207	6.807	6.806	

Notes: [a] Elliott-Rothenberg-Stock (1996, Table 1)
*, †, and ‡ indicate significance at the 1 per cent, 5 per cent, and 10 per cent levels, respectively

[9] Whereas the null was rejected at the 5% level for lnYN and lnREER in Model A that allows for a change in only the intercept.

Table 3: Ng-Perron Test for Unit Root

Null Hypothesis: LNTB has a unit root
Constant

	Level				Differenced				
	MZa	MZt	MSB	MPT	MZa	MZt	MSB	MPT	Decision
lnTB	-8.144†	-2.015†	0.247‡	3.017†	-1.955	-0.869	0.444	11.258	I(0)
lnYN	0.949	0.834	0.878	54.990	-10.183†	-2.236†	0.219‡	2.485†	I(1)
lnYW	1.493	1.597	1.069	86.779	-4.721	-1.325	0.280	5.637	-
lnREER	-2.908	-1.152	0.396	8.293	-38.215*	-4.370*	0.114*	0.642*	I(1)
Asymptotic critical values[b]									
1%	-13.8	-2.58	0.174	1.78	-13.8	-2.58	0.174	1.78	
5%	-8.1	-1.98	0.233	3.17	-8.1	-1.98	0.233	3.17	
10%	-5.7	-1.62	0.275	4.45	-5.7	-1.62	0.275	4.45	

Null Hypothesis: LNTB has a unit root
Constant, Linear Trend

	Level				Differenced				
	MZa	MZt	MSB	MPT	MZa	MZt	MSB	MPT	Decision
lnTB	-12.481	-2.496	0.200	7.308	-6.309	-1.776	0.281	14.441	-
lnYN	-7.079	-1.861	0.262	12.904	-17.425†	-2.948‡	0.169‡	5.251†	I(1)
lnYW	-14.068	-2.643‡	0.187†	6.528‡	-5.309	-1.503	0.283	16.749	I(0)
lnREER	-8.255	-1.971	0.238	11.234	-40.527*	-4.496*	0.110*	2.275*	I(1)
Asymptotic critical values[b]:									
1%	-23.8	-3.42	0.143	4.03	-23.8	-3.42	0.143	4.03	
5%	-17.3	-2.91	0.168	5.48	-17.3	-2.91	0.168	5.48	
10%	-14.2	-2.62	0.185	6.67	-14.2	-2.62	0.185	6.67	

Notes: [b]Ng-Perron (2001, Table 1)
*, †, and ‡ indicate significance at the 1 per cent, 5 per cent, and 10 % levels, respectively

Table 4: Zivot and Andrews Test for Unit Roots with One Structural Break

Null hypothesis: Unit root (non-stationarity) with no structural break

Alternative hypothesis: Structural break in the intercept at unknown breakpoint

MODEL A (Constant)

Variable	LNTB	LNYN	LNYW	LNREER
Break	1986Q2	1991Q2	1997Q1	1986Q1
Y(-1)	-0.119(-3.079)	-0.071(-5.000) †	-0.050(-3.353)	-0.103(-4.865) †
C	-0.004(-0.221)	1.736(5.085)	1.493(3.362)	0.639(4.830)
@TREND	0.0002(0.756)	0.003(4.375)	0.001(3.559)	0.001(1.801)
DU	0.060(2.035)	0.087(5.619)	-0.017(-3.141)	-0.208(-5.039)

Note: Critical Values for the t-statistic (in parenthesis) for Model A are -5.34, -4.8; -4.58 at the 1%(*), 5%(†), and 10%(‡), respectively.

Table 5: Zivot and Andrews Test for Unit Roots with One Structural Break

Null hypothesis: Unit root (non-stationarity) with no structural break

Alternative hypothesis: Structural break in the intercept and the slope coefficient on the trend at an unknown breakpoint

MODEL C (Constant and Trend)

Variable	LNTB	LNYN	LNYW	LNREER
Break	1991Q1	1991Q2	1997Q1	1986Q1
Y(-1)	-0.170(-4.159)	-0.065(-4.364)	-0.049(-3.070)	-0.104(-4.777)
C	-0.089(-2.666)	1.576(4.370)	1.495(3.082)	0.634(4.743)
@TREND	0.006(3.894)	0.003(4.587)	0.0009(3.037)	0.001(0.416)
DU	-0.117(-3.411)	0.075(4.347)	-0.016(-3.091)	-0.217(-3.954)
DT	-0.005(-3.449)	-0.0007(-1.330)	-0.000001(-0.011)	-0.0008(-0.262)

Note: Critical Values for t-statistic (in parenthesis) for Model C are -5.57-5.08 -4.82 at the 1 %(*), 5 %(†), and 10%(‡), respectively. The t-statistic of the coefficient of y_{t-1} determines the stationarity of the series. The unit root [non-stationarity] null hypothesis can be rejected (accepted) if t-statistics is less (more) than the appropriate critical value.

The Gregory-Hansen (GH) Cointegration Test

Following the findings that the data series are by nature, mostly non-stationary stochastic processes, econometric developments regarding the concepts of cointegration are particularly disposed to testing for equilibrium. The Gregory and Hansen (1996) tests for cointegration where the structural break is test-determined and the cointegrating vectors are allowed to change at an unknown time period is employed. The failure to account for breaks can produce misleading results leading to incorrect inference.[10] Therefore, it is necessary to employ non-linear techniques for testing cointegration if the series have structural breaks. Accordingly, the long run properties of the variables in the behavioural equations were examined using the Gregory and Hansen (1996) tests. Table 6 presents the result of the cointegration test. The null hypothesis of no cointegration is rejected (at the five per cent level) using the Level Shift, and Regime Shift formulation type in the ADF* test. Though at the ten per cent level, the formulation type by Level Shift with Trend could also reject the null. In other words, the variables in the static equations are cointegrated. Gregory and Hansen (1996) hinted that rejection by either the standard ADF or the ADF* implies that there is some long run relationship in the data and that if the standard ADF statistic does not reject, but the ADF* does, it implies that structural change in the cointegrating vector may be important. The standard ADF was conducted prior to the GH test in line with Gregory and Hansen (1996) in our study and the results suggest that the null of no cointegration is rejected. Given that both (ADF and ADF*) statistics have rejected the null hypothesis, à la Gregory and Hansen, no inference regarding the occurrence of structural break could be made from the information alone. In this context, it is suggested to determine whether the cointegrating relationship has been subjected to a regime shift. To examine this, we used the Quandt-Andrews unknown breakpoint test (Table A1in the Appendix).[11] The null hypothesis of no

---

[10] Esso (2010) opines that the cointegration framework of Engle and Granger, and Johansen have limitations especially when dealing with economic data containing the structural breaks. In this case, we tend to reject the hypothesis of cointegration, albeit one with stable cointegrating parameters. The reason is that the residuals from cointegrating regressions capture breaks unaccounted for and thus, typically exhibit non-stationary behaviour.

[11] Based on Quandt (1960), Andrews (1993) modifies the Chow test and allows for unknown breakpoints. Basically, the Quandt-Andrews test performs the Chow breakpoint test for every observation over an interval and calculates the supremum of the Fk statistics. Andrews and Ploberger (1994) develop two additional test statistics, the average (ave F) and the exponential (exp F) form. The null hypothesis of no break is rejected if these test statistics are too large.

structural break is rejected by the three test statistics at the five per cent level of significance. The detected date is 1991:Q2.

Table 6: Gregory-Hansen Cointegration Test

Null Hypothesis: No cointegration
Dependent Variable lnTB

MODEL	ADF*	Tb	Zt	Tb	Za	Tb
2: Level Shift	-5.34†	1989Q3	-4.12	1991Q3	-30.57	1991Q3
3: Level Shift with Trend	-5.44‡	1985Q3	-4.09	1984Q2	-30.46	1991Q3
4: Regime Shift	-6.18†	2000Q1	-4.76	1991Q3	-37.36	1991Q3
MODEL	2: Level Shift		3: Level Shift with Trend		4: Regime Shift	
Critical values[c]:	ADF*, Zt	Za	ADF*, Zt	Za	ADF*, Zt	Za
1%	-5.77	-63.64	-6.05	-70.27	-6.51	-80.15
5%	-5.28	-53.58	-5.57	-59.76	-6.00	-68.94
10%	-5.02	-48.65	-5.33	-54.94	-5.75	-63.42

Notes: [c] Gregory and Hansen (1996) Table 1, pp.109
*, †, and ‡ indicate significance at the 1 per cent, 5 per cent, and 10 % levels, respectively

Long Run Estimation

The results of the cointegration test in Table 7 show that long run relationship exists. The OLS estimator is employed to analyse the long run relationship. Two other long run estimators, the fully modified OLS (FMOLS) and dynamic OLS (DOLS), known to provide robust results in small samples sizes, were also used. All the three approaches provide similar results on the long run relations, thus showing the validity of the results. The results of the long run estimation show a significant negative impact of exchange rate shocks on trade balance across the different approaches. In other words, exchange rate depreciation deteriorates trade balance in the long run. A 100 per cent depreciation of exchange rate leads to about 21 per cent deterioration in trade balance. Domestic and world incomes were found to be insignificant, though with different signs; positive for domestic income and negative for world income. Bahmani-Oskooee and Brooks (1999) however argue that the existence of a cointegration

Hansen (1997) derives an algorithm to calculate approximate asymptotic p-values of these tests.

derived from equation (8) does not necessarily imply that the estimated coefficients are stable. These tests also incorporate the short run dynamics into the long run through residuals. The CUSUM and CUSUMSQ statistics are updated recursively and plotted against the break points of the model. The two tests incorporate the short run dynamics to the long run through residuals. The statistics of the two tests are updated recursively and plotted against the break points of the model. Provided that the plot of these statistics falls inside the critical bounds of five per cent significance, one assumes that the coefficients of a given regression are stable. These tests are usually implemented by means of graphical representation.

Table 7: Long Run Estimations of Trade Balance Model.

	OLS	FMOLS (constant)	FMOLS (Linear Trend)	DOLS (constant)	FMOLS (Linear Trend)
Dependent Var:	LNTB	LNTB	LNTB	LNTB	LNTB
LNYN	0.073	0.077	0.023	0.063	-0.017
	(-0.057)	(-0.107)	(-0.135)	(-0.111)	(-0.150)
	[1.279]	[0.714]	[0.170]	[0.571]	[-0.117]
LNYW	-0.107	-0.215	-0.506	-0.054	-0.352
	(-0.256)	(-0.482)	(-0.556)	(-0.498)	(-0.619)
	[-0.420]	[-0.446]	[-0.910]	[-0.108]	[-0.569]
LNREER	-0.209	-0.317	-0.256	-0.208	-0.236
	(0.050)*	(0.095)*	(0.092)*	(0.098)†	(0.102)†
	[-4.119]*	[-3.341]*	[-2.772]*	[-2.119]†	[-2.298]†
C	2.795	6.519	16.203	1.395	12.433
	(-6.515)	(-12.273)	(-16.846)	(-12.674)	(-18.744)
	[0.429]	[0.531]	[0.961]	[0.110]	[0.663]
@TREND	-	-	0.008	-	0.009
			(-0.010)		(-0.011)
			[0.811]		[0.787]
Observations:	123	122	122	122	122
R-squared:	0.4233	0.4097	0.4494	0.4413	0.4522
F-statistic:	29.1143	NA	NA	NA	NA

Notes: Figures in parenthesis (…) and bracket […] are standard error and t-statistic, respectively

*, †, and ‡ indicate significance at the 1 per cent, 5 per cent, and 10 % levels, respectively

Once the long run is established, we proceed in the spirit of Engle and Granger (1987) by specifying the error-correction term to obtain the short run dynamics. A general error correction model (ECM) of equation (8) is formulated as follows:

$$\Delta \ln TB_t = b_0 + \sum_{i=1}^{m} b_1 \Delta \ln TB_{t-i} + \sum_{i=0}^{m} b_2 \Delta \ln YN_{t-i} + \sum_{i=0}^{m} b_3 \Delta \ln YW_{t-i} + \sum_{i=0}^{m} b_4 \Delta \ln REER_{t-i} + \tau ECM_{t-1} + \varepsilon_t \quad (17)$$

Where: τ is the speed of adjustment parameter and ECM is the error correction term, usually the residual obtained from the estimated cointegration model of equation 8. The result is presented in Table A3 in the Appendix and a maximum of two lags was used for the model based on the Schwarz Bayesian Criteria (SBC).

The impacts of REER depreciation, domestic and foreign incomes on trade balance are examined by estimating an error correction model. The result in Table A2 in the Appendix reveals that exchange rate depreciation has a negative significant effect on trade balance in the short run, although, at the ten per cent level. Specifically, a 100 per cent increase in the contemporaneous exchange rate leads to about 13 per cent deterioration in trade balance. The lagged values of trade balance and the contemporaneous domestic income are found to be positively weighty in explaining changes in trade balance in our models. A 100 per cent increase in the immediate past value of trade balance could lead to 59 per cent increase in trade balance in the current period at the one per cent level of significance. In addition, the coefficient of the error correction term, which shows rate of adjustment to equilibrium, is rightly signed and significant. Its value of –0.11 suggests that 11 per cent of the adjustment is achieved every quarter. For J-curve to exist there should be an initial deterioration of trade balance (in the short run) and a subsequent improvement of trade balance (in the long run) from manoeuvred depreciation of the currency. Based on the results, such a phenomenon does not exist in Nigeria. Judging by the (negative) signs from our short run and long run results, one can safely submit that the use of exchange rate adjustment to address trade balance problems could worsen the deterioration. This result is similar to that obtained by Bahmani-Oskooee and Gelan (2012) and Oyinlola *et al.* (2013).

Reliability and Stability Test of the Estimates

In addition, Table A3 in the Appendix presents diagnostic tests of our model and suggests an absence of major diagnostic problems such as serial correlation, non-normality and specification errors. These results indicate that our estimated import price model is well-specified. Thereafter, it is necessary to check for the stability of the import price function. This is because of the importance of the stability of the import price function for an effective trade policy. This therefore makes it necessary to test whether the estimated import price equation has shifted over time as an important part of this empirical study. As can be observed from Figure 1, the CUSUM and CUSUM Square tests of parameter stability indicate that the parameters are stable during the sample period (Figures 1 and 2 in the Appendix), except between 2003 and 2005 in the CUSUM Square plot.

CONCLUSION

This study examines the long run and short run impact of real exchange rate depreciation on trade balance in Nigeria between 1980:Q1-2010:Q4 using the Gregory-Hansen Structural break cointegration test. Our results show the existence of long run relationship among the variables. There is however, no short run impact of exchange rate depreciation on trade balance at the five per cent level; a negative impact was found at the ten per cent level of significance. Thus, we find no support for J-curve in Nigeria, rather, the trade balance seems to deteriorate both in the short run and long run. A number of policy issues are discerned from the results. The need for further investigation would be insightful especially to understand the determinants of the components of trade balance. Findings from such enquiries could provide support for or caution on the use of exchange rate policy as instrument of adjustment of trade balance.

REFERENCES

Akbostanci, E. 2004. Dynamics of the trade balance: The Turkish J-curve, *Finance and Trade,* 40 (5), 57-73.

Andrews, D. W. K. 1993. Tests for parameter instability and structural change with unknown change point. *Econometrica*, 61(4):821-56.

Andrews, D. W. K. and Ploberger, W. 1994. Optimal tests when a nuisance parameter is present only under the alternative. *Econometrica*, 62(6):1383-1414.

Awan, R., Shahbaz, M. and Javed, K. 2012. Does J-curve phenomenon exist in Pakistan? A revisit. *Interdisciplinary Journal of Contemporary Research in Business*, Vol. 3, NO 9

Aziz, J. 2008. Real and financial sector linkages in China and India. *IMF Working Paper 08/95.* International Monetary Fund: Washington, DC.

Baharumshah, A. Z. 2001. The effect of exchange rate on bilateral trade balance: New evidence from Malaysia and Thailand. *Asian Economic Journal*, Vol. 15, No. 3,291 - 311

Bahmani-Oskooee, M. and Brooks, T.J. 1999. Bilateral J-curve between US and her trading partners. *Weltwirtschaftliches Archiv*, 135, 156-165.

Bahmani-Oskooee, M. 1985. Devaluation and the J-curve: Some evidence from LCDs. *The Review of Economics and Statistics*, 67(3), 500-4

Bahmani-Oskooee, M. and Goswami, G. G. 2003. A disaggregated approach to test the J-curve phenomenon: Japan versus her major trading partners. *International Journal of Economics and Finance*, 27, 102-113.

Bahmani-Oskooee, M. and Tatchawan, K. 2001. Bilateral J-curve between trading partners. *Journal of Economic Development* , 107-117.

Bahmani-Oskooee, M. and Ratha, A. 2004a The J-curve: A literature review. *Applied Economics*, 36, 1377-1398.

Bahmani-Oskooee, M. and Ratha, A. 2004b. The J-curve dynamics of US bilateral trade. *Journal of Economics and Finance*, 28, 32-38.

Bahmani-Oskooee, M., Economidou, C. and Goswami, G. G. 2006. Bilateral J-curve between the UK vis-à-vis her major trading partners, *Applied Economics*, 38, 879-888.

Bahmani-Oskooee, M. and Gelan, A. 2009. How stable is the demand for money in African countries. *Journal of Economic Studies*, 36, 216-235.

Carter, C. and Pick, D. 1989. The J-curve effect and the US agricultural trade balance. *American Journal of Agricultural Economics,* 71 (3), 712-720.

Davinder, K.S. and Shome, S. 2013. Evidence of a J-curve in India. *International Monthly Refereed Journal of Research in Management and Technology*, 2, 15-23.

DeJong, D., Nankervis, J. C., Savin, N. E. and Whiteman, C. H. 1992. The power problems of unit root test in time series with autoregression errors. *Journal of Econometrics* , 53(1-3),323-343.

Dickey, D.A. and Fuller, W. A. 1979. Distribution of the estimators for autoregressive time series with a unit root. *Journal of the American Statistical Association*, 74, 427-431.

Elliott, G., Rothenberg, T. J. and Stock, J. H. 1996. Efficient tests for an autoregressive unit root. *Econometrica,* 64, 813-836.

Engle, R. and Granger, C.W.J. 1987. Cointegrtaion and error correction: representation, estimation and testing, *Econometrica*, 55, 251-276.

Esso, L.J. 2010. Re-examining the finance-growth nexus: Structural break, threshold cointegration and causality evidence from the ECOWAS. *Journal of Economic Development*, 35(3), 57-79.

Glynn, J., Perera, N. and Verma, R. 2007. Unit root tests and structural breaks: A survey with spplications. *Revista de Metodos Cuantitativos para La Econom´ia y la Empresa,* (3). Pp. 63–79. June.

Gregory, A. W. and Hansen, B. E. 1996. Residual-based tests for cointegration in models with regime shifts, *Journal of Econometrics*, 70, 99–126.

Gregory, A.W., Nason, J.M., Watt, D. 1996. Testing for structural breaks in cointegrated relationship. *Journal of Econometrics*, 71, 321-341.

Gupta-Kapoor, A. and Ramakrishna, U. 1999. Is there a J-curve? A new estimation for Japan. *International Economic Journal,* 13 (4), 71-19.

Hacker, R.S. and Hatemi, A.J. 2004. The effect of exchange rate changes on trade balances in the short and long run. *Economics of Transition*, 12, 777-799.

Hansen, B. E. 1997. Approximate asymptotic p values for structural-change tests. *Journal of Business & Economic Statistics*, 15(1):60-67.

Hsing, H. 2005. Re-examination of J-curve effect for Japan, Korea and Taiwan. *Japan and the World economy,* 2 (17), 43-58.

Khatoon, R. and Rahman, M. 2009. Assessing the existence of the J-curve effect in Bangladesh. *The Bangladesh Development Studies*, Vol. XXXII, June 2009, No. 2

Koray F. and McMillin, W. D. 1999. Monetary shocks, the exchange rate, and the trade balance. *Journal of International Money and Finance,* 18, 925-940.

Krugman, P.R. and Baldwin, R. E. 1987. The persistence of the US trade deficit. *Brookings Papers on Economic Activity,* 1, 1-43.18

Krugman, P. and Obstfeld, M. 2001. *International economics: Theory and policy*, 5th ed., Addison-Wesley, New York.

Leonard, G. and Stockman, A. C. 2001. Current accounts and exchange rates: A new look at the evidence. *NBER Working Paper,* No: 8361.

Magee, S.P. 1973. Currency contracts, pass-through, and devaluation. *Brookings Papers of Economic Activity, 1*: 303-323.

Meltzer, L. 1948. The theory of international trade. In: H. S. Ellis (Ed), *A survey of contemporary economics.* Philadelphia: Blakiston.

Narayan, P. K. 2004. Reformulating critical values for the bounds F -statistics approach to cointegration: An application to the tourism demand model for Fiji. *Department of Economics Discussion Papers No. 02/04*, Monash University, Melbourne, Australia.

Narayan, P.K. and Narayan, S. 2004. The J-curve: Evidence from Fiji. *International Journal of Applied Economics*, 18, 369-380.

Nasir, A., Chowdhury, A. and Mridha, H. 2009. Do trade balances adjust to exchange rate shocks? Evidence from bilateral trade of Bangladesh with G-5 countries. *International Journal of Applied Economics*, 6(2), September 2009, 51-69

Ng, S. and Perron, P. 2001. Lag length selection and the construction of unit root tests with good size and power. *Econometrica*, 69, 6, 1519-1554

Onofowora, O. 2003. Exchange rate and trade balance in East Asia: Is there a J-curve? *Economics Bulletin*, 5(18), 1-13.

Oyinlola, M. A., Omisakin, O.A. and Adeniyi, O.A. 2013. Do trade balances adjust to exchange rate shocks? An enquiry evidence of J-curve in Nigeria. *The Social Sciences, 8: 72-79*

Perron, P. 1997. Further evidence on breaking trend functions in macroeconomic variables. *Journal of Econometrics*, 80, 355-385.

Perron, P. 1989. The great crash, the oil price shock and the unit root hypothesis. *Econometrica,* 57, 1361-1401.

Phillips, P.C. B. and Perron, P. 1988. Testing for a unit root in time series regression. *Biometrika*, 75, 335-359.

Quandt, R. E. 1960. Tests of the hypothesis that a linear regression system obeys two separate regimes. *Journal of the American Statistical Association*, 55(290):324-330.

Robinson, J. 1947. *Essays in the theory of employment.* Second ed., Oxford: Basil Blackwell.

Rose, A.K. and Yellen, J.L. 1989. Is there a J-curve? *Journal of Monetary Economics*, 24, 53-68.

Said E. Said and David A. Dickey, 1984. Testing for unit roots in autoregressive {Moving average models of unknown order. *Biometrika*, 71:599-607.

Schwert, G.W. 1989. Tests for unit roots: A Monte Carlo investigation. *Journal of Business and Economic Statistics*, 7, 147-160.

Simakova, J. 2012. Bilateral J-curve between Slovakia and its major trading partners. Proceedings of 30th International Conference Mathematical Methods in Economics.

Singh, T. 2002. India's trade balance: The role of income and exchange rates. *Journal of Policy Modelling,*

Wilson, P. 2001. Exchange rates and the trade balance for dynamic Asian economies: Does the J-curve exist for Singapore, Malaysia and Korea? *Open Economies Review*, 12, 389-413.

Zivot, E. and Andrews, K. 1992. Further evidence on the great crash, the oil price shock, and the unit root hypothesis. *Journal of Business and Economic Statistics*, 10 (10), pp. 251–70.

APPENDIX

Figure 1A: Nigeria Exports: Jan. 2002-Dec. 2012

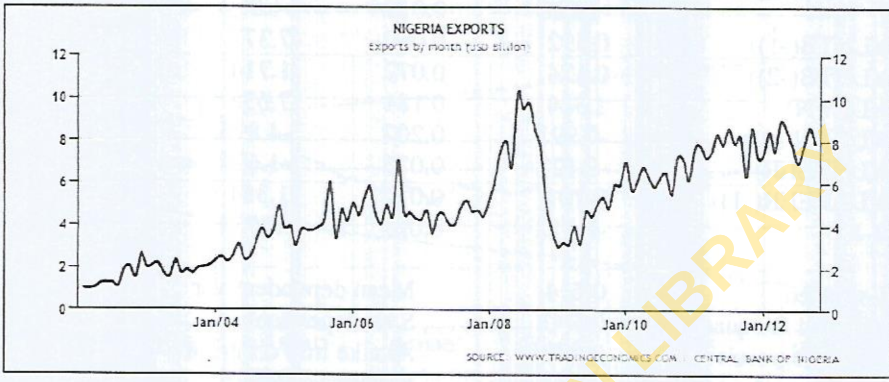


Figure 2A: Nigeria Imports: Jan. 2002-Dec. 2012

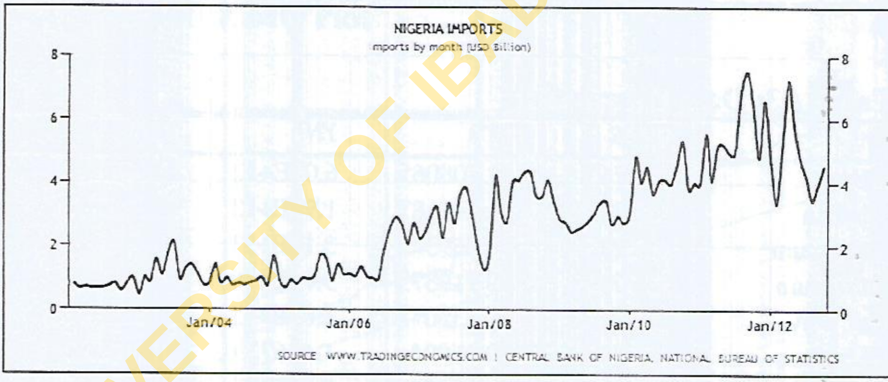


Table A1: Quandt-Andrews unknown breakpoint test

Null Hypothesis: No breakpoints within 15% trimmed data

Statistic	Max.LR F-statistic (1991Q2)	Max. Wald F-statistic (1991Q2)	Exp LR F-statistic	Exp Wald F-statistic	Ave LR F-statistic	Ave Wald F-statistic
Value	20.30	81.19	6.58	36.40	7.58	30.30
Prob.	0.01	0.00	0.01	0.00	0.05	0.00

Note: P-values are calculated according to Hansen (1997).

Table A2: Parsimonious Short Run Dynamics of Trade Balance

Variable	*Coefficient*	*Std. Error*	*t-Statistic*	*Prob.*
C	-0.020	0.010	-1.919	0.058
D(LNTB(-1))	0.592	0.080	7.377	0.000
D(LNTB(-2))	0.124	0.072	1.714	0.089
D(LNYN)	1.384	0.181	7.636	0.000
D(LNYN(-1))	-0.991	0.202	-4.896	0.000
D(LNREER)	-0.127	0.075	-1.697	0.093
D(LNREER(-1))	0.101	0.074	1.364	0.175
E(-1)	-0.112	0.029	-3.934	0.000
R-squared	0.614	Mean dependent var		0.002
Adjusted R-squared	0.590	S.D. dependent var		0.110
S.E. of regression	0.070	Akaike info criterion		-2.409
Sum squared resid	0.553	Schwarz criterion		-2.223
Log likelihood	152.517	F-statistic		25.440
Durbin-Watson stat	2.107	Prob(F-statistic)		0.000

Table A3: Data Description

	REER	*TB*	*YN*	*YW*
Mean	176.75	1.708065	6.09E+12	2.97E+13
Median	106.0048	1.612187	1.97E+12	2.97E+13
Maximum	652.9956	2.769375	3.14E+13	6.66E+13
Minimum	53.97875	0.719375	5.02E+10	1.07E+13
Std. Dev.	156.3635	0.538013	8.64E+12	1.53E+13
Skewness	1.71422	0.16994	1.446739	0.702082
Kurtosis	4.740036	2.098115	3.745307	2.585551
Jarque-Bera	76.37327	4.799388	46.12644	11.07445
Probability	0	0.090746	0	0.003937
Sum	21917	211.8	7.55E+14	3.68E+15
Sum Sq. Dev.	3007294	35.60329	9.19E+27	2.89E+28
Observations	124	124	124	124

Figure 1: CUSUM Plot

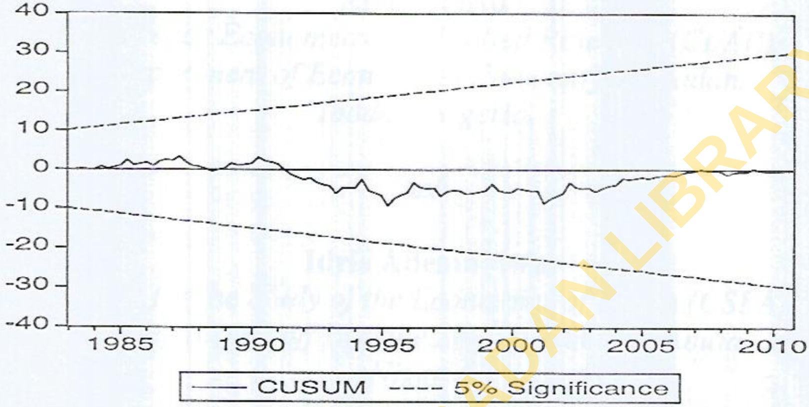


Figure 2: CUSUMQ Plot