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GENERAL & APPLIED ECONOMICS | RESEARCH ARTICLE

Fiscal sustainability in the presence of structural breaks: Does overconfidence on resource exports hurt government's ability to finance debt? Evidence from Nigeria

Aliyu Alhaji Jibrilla^{1*}

Abstract: The sustainability of the Nigerian fiscal deficit along with the role of the dynamics of government revenues and spending in adjusting the size of the deficit is examined using annual data from 1961 to 2014. After allowing for structural breaks, the study finds evidence of a cointegration relation between the government revenues and spending. The results did not indicate the presence of asymmetries in either the threshold autoregression or momentum threshold autoregression specifications of the country's budgetary adjustment process. Interestingly, the size of the long run slope parameter appears to be significantly less than one, thus offering support for the soft budget strategy, which also suggests that the government might face difficulties in financing its debts in the long run. Lastly, the short run and long run Granger causality results, while providing evidence in support of the fiscal synchronization hypothesis, also raise some important issues, particularly on the strength of budget deficit sustainability in the country.

Subjects: Economics; Economics and Development; Finance; Sustainable Development

Keywords: revenues; expenditures; oil price volatility; structural breaks; budget deficit sustainability; causality; error correction

JEL Classifications: C1; E62; H6; Q32

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PUBLIC INTEREST STATEMENT

Crude oil price volatility in the global market tends to have a significant influence on the revenue flows and expenditures in a country that typically depends on oil such as Nigeria. In such a country, policy-makers are confronted with a challenge of setting policies that will restrain the possible negative spillover effect of oil price volatility on the size of the federal budget deficit. This paper evaluates how government revenues and spending adjust to contain the size of the budget deficit in Nigeria as well as the ability of the government to market its future debt. The empirical findings suggest that over-dependence on oil exports tends to threaten the country's ability to redeem its debts from bondholders at maturity without difficulties. Thus, the paper calls on policy-makers to devise effective policies that will channel revenue windfall from crude oil sales into infrastructure development to make the economy more diversified.

1. Introduction

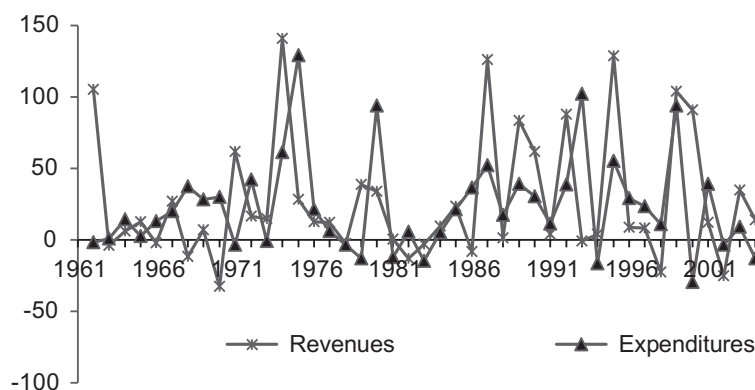
The recent sharp fall in the global crude oil prices which began in late 2014 has renewed concerns about the fiscal sustainability of oil-dependent economies. Nigeria is one of the oil-exporting countries that is frequently facing the problem of budget deficit due in large part to volatility in the international oil prices. Overconfidence on oil export revenues—belief that the windfall from crude oil sales will last for an extended period—has typically led to a failure by successive governments in Nigeria to adopt sufficiently prudent expenditure policies in good times (see, e.g. Budina, Pang, & Van Wijnbergen, 2007). Besides, history has shown that global oil prices have never been regarded to be stable (Dibra, 2015). Instability of fiscal revenues will likely make effective planning and management of government budget tough. In instances of oil price increases, higher revenue growth could create unrealistic anticipations about the amount of public outlays that should be sustained over time. When there is a subsequent downturn in prices of crude oil, it may not be followed by necessary adjustment of government expenditures and revenues to counterbalance the massive budget deficits that can unavoidably emerge. The likely effect will be an undesirable fiscal consequence over the long term, including structural budget deficits and rising debt. For instance, during the dramatic increase in oil prices in 1973–1974, the higher revenue streams induced the government to raise public spending (Adedipe, 2004).

However, the shortfall in revenues due to falling crude oil prices in the international markets (during 1975–1979) forced the government to resort to (massive) public borrowing to finance the (budget deficit) development programs and projects of the third national development plan. There was also a similar incidence of increased public borrowing to fund federal budget deficit during the fourth national development plan (see Bienen, 1983; Budina et al., 2007; Lewis, 1994; Oladipo, 2013). In fact, the levels of public debt in the country remain high for an extended period thanks to the “debt pardon in 2005” (see Budina et al., 2007). Could this be an indication that the Nigerian Government faces difficulties in financing its future debt? Answering this question requires, at least, information on the relationships between government revenues and expenditures and how they adjust to contain fiscal deficits in the country.

Understanding both short run and long run behavior of government revenue and expenditure is central to the prudent budgetary process and fiscal policy sustainability. A sustainable fiscal policy may refer to “policy that can be pursued, however long without any major interventions in tax and spending patterns” (Krejdí, 2006, p. 4).

Fiscal policy in Nigeria has been heavily influenced by shocks in the international oil prices, as noted above, which reflects two challenges that face budget planning and administration in the country. First, to ensure that, the expected and desired long run impact of revenues and expenditures on the national economy is compatible with the sustainable use of oil resources. Second, to restrain the short run spillover effect of income volatility that may follow oil price shock(s) into the national budget (see Baunsgaard, 2003). Figure 1 plots the percentage of variation in public

Figure 1. Percentage of variation in public revenues or expenditures as % of GDP, 1961–2014.



revenues or spending as a share of real GDP from 1961 to 2014. Although the government revenues had increased by 105% in 1962 from 1961, it has however fallen to less than 29% and -1.76% in 1975 and 2014, respectively. In contrast, government spending increased from a low of -1.84% in 1962 (from 1961) to a high of more than 128% in 1975, falling back to -16% in 2014.

The variation in public revenues has largely been due to a volatility in the international oil prices and undiversified export base over the whole period (see, e.g. Aizenman & Pinto, 2005; Budina & van Wijnbergen, 2008; Cottarelli, 2012; Ghura & Pattillo, 2012; Jibrilla & Mohammed, 2015 among others). However, the variations in public expenditures cannot be explained solely by crude oil price volatility; it has also been determined by government borrowing to finance development projects. For example, the higher percentage variation in public spending that reached more than 128% in 1975 was due to massive loans from domestic and multilateral sources, as noted above. Higher public debts are critical to a nation's ability to sustain fiscal deficit in the long run. Fiscal sustainability requires the government to be solvent—be able to pay its debt at a future date. But a highly indebted country will need to devote a substantial share of national income to service the accumulated debts with the consequences of whether the government should resort to money creation/seigniorage, debt rescheduling, be forced to seek for debt forgiveness, surprise inflation, or default (see Bleaney, 1996; Labonte, 2012). For example, the increasing size of public borrowing over the past few years had resulted in a high debt service burden that compelled the government to seek for debt relief in the later part of 2004 from the Paris Club, which they granted in June 2005 (Alsop & Rogger, 2008; Dijkstra, 2013).

The lower percentage variation (Figure 1), which has fallen to as low as -16% in 2014 from 2013, was due to the increasing debt profile of the country, a situation that forced the government to reduce capital expenditure by more than 30% from 2013, and increase debt service spending by more than 20% compared to less than 6% in 2013 (PWC, 2014). This situation tends to hurt public investment in critical infrastructures, dearly needed by the country for sustainable growth. In fact, increasing public borrowing, whenever there are episodes of declining revenue, has been the experience of Nigeria since the discoveries of oil in the early 1970s (see, e.g. Budina et al., 2007; Oladipo, 2013).

Although public borrowing is essentially vital for an economy to function, especially when there is revenue shortfall, high levels of debt could trigger rising inflation which, in turn, could encourage capital flight. The resulting effect of these outcomes might jeopardize the government's ability to control its budget priorities. Besides, curtailing government capital spending to service debt could reduce improvement in long-term fiscal balance or even a worsening (see, e.g. Bleaney, Gemmell, & Greenaway, 1995; Bruce & Turnovsky, 1999 among others).

While some degree of instability in the level and composition of government spending may be inevitable in Nigeria due to the historical linkage of its budgets to natural resource exports, the lack of diversified export base, which implies limited tax base tended to question the ability of the government to balance its budget in the long run. This study, therefore, intends to examine the validity of this claim.

Although, at a theoretical level, there is no precise standard for measuring fiscal sustainability; consistency of the budgetary process may be an indication that a country's budget is sustainable. However, Quintos (1995), Cunado, Gil-Alana, and Pérez de Gracia (2004), and Payne, Mohammadi, and Cak (2008), among others, note that cointegration between public revenues and expenditures may indicate an evidence of inter-temporal equilibrium condition or budget constraint, which may also suggest fiscal sustainability.

The degree of this sustainability can further be examined within the framework of the soft budget constraint (SBC) strategy and/or hard budget constraint (HBC) policy (see, e.g. Paleologou, 2013). The SBC phenomenon occurs if, for example, the federal government is ever ready to cover all or part of any state government's budget deficit that may arise due to its inability to cover its spending out of

its revenues/taxes. In contrast, the HBC phenomenon occurs if such a bailout package seldom exists (see e.g. Kornai, Maskin, & Roland, 2003; Rodden, Eskeland, & Litvack, 2003).

Each of the two alternative strategies offers a prospect of being able to explain budgetary behavior in Nigeria. For example, it is not unusual for subnational governments that have become insolvent and face a revenue crisis to apply for credit rescue package (typically short-term trade credits), and usually with absolute guarantees by the federal government through the export credit agency or even for the federal bailout fund (Budina et al., 2007; CBN, 2015). It is also a common practice in the country that during “good times” the federal government often shares revenue surplus, often referred to as “excess crude” with the states and local governments. Since this has a tendency to facilitate an expansion of spending programs at both the national and subnational governments, it may further restrain the ability of the government to smooth expenditure, with likely implications for highly volatile exchange rates, rising interest rate, and slow economic growth (see Baunsgaard, 2003; Cunado et al., 2004). However, the government has recently taken steps to control unsustainable expenditure and improve efficiency. As a result, HBCs have been imposed on federal parastatals (see, e.g. Nowak & Muñiz, 2004).

This study, therefore, intends to answer the following questions. Is there cointegration between government revenues and expenditures in Nigeria? How is the relationship affecting the sustainability of budget deficits? Which strategy dominates fiscal policy in the country? In other words, does the Nigerian Government run the economy on SBC or HBC strategy? What roles do government revenues and expenditures play in the adjustment of both short-term and long-term budget deficits in the country?

The main contributions of this study are twofold. First, evaluating the adjustment behavior of the government revenues and spending due to budgetary disequilibria build up the contributions of this study to Nigeria’s fiscal sustainability. Investigating this issue is particularly important for policy-makers considering that fiscal policy in the country has been influenced heavily by volatility in the international oil prices (see e.g. Baunsgaard, 2003). The most recent studies that examined this issue in the country include the works of Obioma and Ozughalu (2010) and Nwosu and Okafor (2014). The primary tool used to analyze the sustainability of government budget deficits in these studies is cointegration tests between public revenues and expenditures. The determination of the degree of fiscal sustainability is important from the perspective of fiscal policy-making in Nigeria as it can better inform policy-makers to choose appropriate intervention measures in the design of effective fiscal rule. This issue has however been largely ignored in the country’s literature. Quintos (1995) shows that if public revenues and spending are cointegrated, and the coefficient relating them is significantly equal to one, it is an indication that the budget deficit is “strongly” sustainable. However, if the coefficient is significantly less than one, then it (the deficit) can be considered “weakly” sustainable.

Second, a common argument in the empirical economic literature is that the standard cointegration tests tend to have limited power when variables are exposed to exogenous or endogenous breaks (Gregory & Hansen, 1996; Johansen, Mosconi, & Nielsen, 2000). This issue appears particularly relevant to Nigeria, given that budgeting in the country largely depends upon crude oil exports; shocks in the global oil prices may lead to structural shifts which can affect the long run relationship between revenues and spending. Accordingly, the present study complements the analysis of long run fiscal policy sustainability by accounting for both exogenous and endogenous shifts that might affect the budgetary process in the country between sample periods, an issue that has largely been overlooked, or assumed irrelevant by researchers within the country’s fiscal policy literature. The rest of the paper is organized as follows: Section 2 reviewed related literature and discusses the theoretical models; Section 3 describes empirical methodology and data; Section 4 presents empirical findings; and; finally, Section 5 concludes.

2. Literature review

Theoretically, there have been four hypotheses on the government revenues–government expenditures literature: tax/revenue-and-spend hypothesis, spend-and-tax/revenue hypothesis, the fiscal synchronization hypothesis, and institutional separation hypothesis. Under the tax/revenue-and-spend hypothesis, changes in government revenues cause the government spending to change (Friedman, 1978). Under the spend-and-tax/revenue hypothesis, expenditure decisions are first set by the government and then followed by required adjustments in tax policy and revenues to accommodate such expenditure. According to this hypothesis, higher government spending leads to higher taxes/revenues (Peacock & Wiseman, 1979). “This can be achieved through both direct and indirect taxation, borrowing, and money creation” (Kollias & Makrydakis, 2000, p. 536). Under the fiscal synchronization hypothesis, the government expenditure decisions and that of revenue necessary to finance such spending are made simultaneously (see, e.g. Meltzer & Richard, 1981). The institutional separation hypothesis suggests an independent relationship between expenditure decisions and taxing decisions (see Payne, 2003; Peacock & Wiseman, 1979; Roberts, 1978).

The empirical testing of these hypotheses, as noted above, has been built on a cointegration relationship between government revenues and expenditures. For example, the tax/revenue-and-spend hypothesis, which suggests that increases in public taxes or revenues cause government spending to increase, suggests that benefits from such funds will not likely reduce federal budget deficits. Mounts and Sowell (1997) using the US data, Narayan and Narayan (2006) for developed and developing countries, including Chile, El Salvador, Haiti, Mauritius, and Venezuela, Eita and Mbazima (2008) for Namibia, Payne et al. (2008) for the case of Turkey, Obioma and Ozughalu (2010) using Nigerian data, Westerlund, Mahdavi, and Firoozi (2011) for local government units in the US, and Apergis, Payne, and Saunoris (2012) for the case of Greece, among others, have empirically validated the tax/revenue-and-spend hypothesis.

The spend-and-tax/revenue hypothesis, which predicts that temporary increases in government expenditures will lead to permanent increases in taxes/revenues, implies a positive unidirectional causality from government spending to taxes/revenues. Using data from OECD countries, Joulfaian and Mookerjee (1990) provide empirical evidence for the spend-and-tax hypothesis. A similar finding was offered by Jones and Joulfaian (1992) for the US, and recently Saunoris and Payne (2010) using UK data, and Paleologou (2013) with asymmetric adjustment toward the long run equilibrium for Greece, among others.

The fiscal synchronization hypothesis, which assumes that government revenues and spending adjust contemporaneously, points to a bidirectional causality between taxes/income and spending. Kollias and Makrydakis (2000) find evidence for the fiscal synchronization hypothesis using data from Greece and Ireland, Li (2001) offered similar evidence for China, Nyamongo, Sichei, and Schoeman (2007) for South Africa, and Paleologou (2013) for Sweden and Germany, among others. Lastly, empirical evidence for the institutional separation hypothesis, which asserts an independent relationship between government revenues and spending, is provided by Hoover and Sheffrin (1992) for the US after accounting for breaks in the data-set; similar evidence for the US was also found by Baghestani and McNown (1994), and Narayan and Narayan (2006) for the countries which include Ecuador, Guatemala, and Uruguay, among others.

In a study that linked the four afore-discussed hypotheses with the concepts of HBC and SBC, Paleologou (2013) argued that the four alternative hypotheses are not equally feasible under both the two budget constraint strategies. He mainly contended that while the fiscal synchronization and tax-and-spend hypotheses would possibly hold for an economy that runs under a policy of HBC, the spend-and-tax hypothesis would likely hold for a government that runs its economy on a SBC strategy. The author offers evidence in support of the SBC strategy for Greece after allowing for a nonlinear adjustment in the long run relationship between revenues and expenditures.

The question of whether government budget deficit is sustainable has also been investigated within the concept of strong or weak sustainability (Quintos, 1995). The decision on the strength of sustainability is made based on the outcome of the coefficient relating government revenues and spending. If the estimated (slope) coefficient is not significantly different from one, it is an indication that “strong” sustainability exists, while a “weak” sustainability can be said to exist when the coefficient is significantly less than one.

An early empirical test of this concept which provides evidence for sustainability was offered in a study by Quintos (1995) for the US. A later study Cunado et al. (2004) also used a similar concept and found a case of weak sustainability of US deficit when a structural break in the mid-1970s was allowed. A recent study by Payne et al. (2008) provides evidence for weak sustainability using Turkey’s data. Similar evidence was found in a more recent work by Paleologou (2013) for Greece, as mentioned above. However, best known, comparable evidence on the issue of fiscal sustainability remains less well known in the case of Nigeria.

Very limited empirical investigations have been done to examine the long run relationship between revenues and spending in Nigeria in the context of the afore-discussed hypotheses, yet no definite conclusion could be reached from the available evidence. Studies that examined these hypotheses in the country include the work of Obioma and Ozughalu (2010) who found cointegration between revenues and expenditures and provided supportive evidence for the tax-and-spend hypothesis. This finding suggests that an increased revenue induces the government to expand its spending activities, which seems unlikely to reduce fiscal deficits. Testing the same assumptions, as in Obioma and Ozughalu (2010), Nwosu and Okafor (2014) found contrary evidence (though they also found cointegration between the two variables), which suggests that government expenditure instigates revenues in Nigeria. A study by Aregbeyen and Ibrahim (2012), however, could not reject the null hypothesis of no cointegration when they used revenues as the dependent variable, but found evidence that supported fiscal synchronization hypothesis.

Despite the findings by Obioma and Ozughalu (2010), Nwosu and Okafor (2014) provide evidence that indicates fiscal deficit in Nigeria is sustainable; their studies are however limited because they ignored to examine further the degree of that sustainability. Limited empirical studies on the degree of budget deficit sustainability, particularly from Nigeria, also meant the country could not take advantage of factual evidence for the evaluation of policy options needed to design effective fiscal rule in the country. In subsequent sections, this issue is addressed.

2.1. Theoretical models: government budget constraint

To examine the relationship between government revenues and expenditures in Nigeria, this study adopts the common budget deficits sustainability framework, which incorporates essential factors that influence the government’s one-period budget constraint. Formally, the model can be represented as (Barro, 1974; Paleologou, 2013; Payne et al., 2008; Quintos, 1995)

$$GG_t + (1 + i_t)B_{t-1} = R_t + B_t \quad (1)$$

where GG_t is the purchases of goods, services, and transfers by the government; B_t denotes the value of government debt; R_t represents government taxes/revenues; and i_t is the real interest rate.

Note that Equation (1) can be rewritten as

$$B_{t-1} = (R_t - GG_t)(1 + i_t)^{-1} + B_t(1 + i_t)^{-1} \quad (2)$$

Substituting for B_t yields

$$B_0 = \sum_{t=1}^{\infty} r_t(R_t - GG_t) \quad (3)$$

where $r_t = \prod_{s=1}^t \sigma_s$ and $\sigma_s = (1 + i_s)^{-1}$. The inter-temporal budget solvency requires that future surpluses must finance the current debt, which also implies that the stock of government debt should not, on average, grow faster than the rate of economic growth. This further suggests that the limit of the government debt term in Equation (3) is equal to zero, $\lim_{n \rightarrow \infty} r_n B_n = 0$. In the absence of this condition, the government is allowed to undertake a Ponzi scheme in which the new debt should finance the maturing one. It follows that the government's budget constraint can be written as

$$B_0 = \sum_{t=1}^{\infty} r_t (R_t - GG_t) + \lim_{n \rightarrow \infty} r_n B_n \quad (4)$$

Assuming that the real interest rate i_t is stationary with an unconditional mean (Hakkio & Rush, 1991), Equation (1) may be expressed as

$$R_t = \theta_0 + \psi_1 G_t + \varepsilon_t \quad (5)$$

where R_t is the logarithmic of government taxes/revenues; G_t is the logarithmic of government spending that includes an interest payment on its debt; θ_0 and ψ_1 denote the cointegrating parameters; and ε_t is the error term that may be serially correlated, which reflects the budgetary disequilibrium between government revenues and its spending. Thus, fiscal sustainability (or sustainability of the budget deficit) requires cointegration between R_t and G_t . It has been argued that the strength of fiscal sustainability depends on the estimated value of ψ_1 . For example, on the one hand, if R_t and G_t are cointegrated and $\hat{\psi}_1 = 1$, then the budget deficit exhibits "strong" sustainability, and/or it means that the government follows a HBC strategy; on the other hand, if R_t and G_t are cointegrated and $0 < \hat{\psi}_1 < 1$, the deficit exhibits "weak" sustainability, which further indicates the government follows a SBC strategy (see, for instance, Paleologou, 2013; Payne et al., 2008). Moreover, "strong" fiscal sustainability also requires that the budget deficit (BD) is a stationary process (see Payne et al., 2008). In other words, sustainability requires the budget deficit variable has to be stationary or integrated of order zero, $I(0)$.

However, if the budget deficit is found to be unit root or, for example, integrated of order one, $I(1)$, it is an indication of a "weak" sustainability, thereby questioning the ability of the government to finance its future debt in the long run. For example, if government spending is more than its revenues, it increases the risk of default, making convergence of the debt term in Equation (4) to zero very difficult.

3. Empirical methodology and data

3.1. Unit root tests

As budget sustainability is a long run equilibrium concept, this study used Equation (5) to estimate the long run relationship between revenues and spending in Nigeria. Before proceeding to the identification of a possible long run relationship among the study variables, there is a need to determine their order of integration using Augmented Dickey-Fuller (ADF, Dickey & Fuller, 1979), the Phillips-Perron (1988) (PP), and Kwiatkowski-Phillips-Schmidt-Shin (KPSS) unit root tests. The null hypothesis under the ADF and PP tests is that the observed variable (tested) has a unit root, while under the KPSS test, the variable is stationary around a deterministic trend. Once the variables are found to be stationary, the next step will involve examining the long run relationship between government revenues and spending.

It is worth mentioning that because revenues and expenditures in Nigeria largely depend upon the earnings from the oil exports, fluctuations/shocks in the international oil prices may lead to structural shifts which can affect the integration order of the variables over the sample period. If the structural breaks exist, it will influence the inference in the unit root tests conducted using the ADF, PP, and KPSS (see Perron, 1989). To account for possible structural breaks, the analysis employs a

methodology proposed by Zivot and Andrews (ZA, 1992) that endogenously corrects for one structural break and technique by Lumsdaine and Papell (LP, 1997) for two structural breaks in the data-set.

To allow for more general specifications, a model proposed by ZA (1992) that allows for a shift in both the intercept and slope is considered to examine the unit root properties of revenues and expenditures. The model has the following form:

$$\Delta y_t = k + \alpha y_{t-1} + \beta t + \theta_1 DU_t + \gamma_1 DT_t + \sum_{j=1}^k d_j \Delta y_{t-j} + \varepsilon_t \quad (6)$$

where Δ is the first difference operator, t represents time in years, $t = 1, 2, 3, \dots, T$. The break date is denoted by TB_t . DU_t is a dummy indicator variable for a mean shift occurring at times TB_t and DT_t for $DU_t = 1$ and $DT = t - TB_t$ if $t > TB_t$; 0 otherwise (see Narayan, 2005). The terms Δy_{t-j} in Equation (6) account for possible serial correlation problems and ensure that the errors are white noise. The model tests the null hypothesis that the series, y_t , is an integrated process without a structural break, against a one-break alternative in the trend function which occurs at an unknown period. The breakpoint is determined at the minimum value of the ADF statistic within [.15, .85] “trimming region.” The critical values of this test are based on ZA (1992). Although ZA (1992) is used extensively in the applied economic literature (see, e.g. Chang, 2002; Chow, Cotsomitis, & Kwan, 2002; Narayan, 2005; Worthington & Pahlavani, 2007 among others), it has been argued that the unit root test model that allows for only one structural break may produce results that can lead to erroneous conclusions when such a model is estimated on data that are characterized by two breaks.

Lumsdaine and Papell (1997) extend the ZA (1992) method to allow for two endogenous structural breaks as follows:

$$\Delta y_t = k + \alpha y_{t-1} + \beta t + \theta DU1_t + \gamma DT1_t + \omega DU2_t + \psi DT2_t + \sum_{j=1}^k d_j \Delta y_{t-j} + \varepsilon_t \quad (7)$$

Table 1. Descriptive statistics of the variables for the period 1981–2011

| | Mean | Median | Min. | Max. | Std. Deviation |
|-------|-------|--------|--------|-------|----------------|
| R_t | .0037 | .0001 | -.0001 | .0235 | .0062 |
| G_t | .0018 | .0001 | -.0001 | .0102 | .0028 |

Note: Figures are in billion naira.

Table 2. Results of the ADF, PP, and KPSS unit root tests

| Variables | ADF | PP | KPSS |
|---------------|-----------|------------|---------|
| R_t | -2.290 | -2.175 | .225*** |
| ΔR_t | -7.699*** | -7.912*** | .252 |
| G_t | -2.812 | -2.812 | .214*** |
| ΔG_t | -8.259*** | -8.874*** | .205 |
| BD_t | -4.109** | -4.109** | .224*** |
| ΔBD_t | -7.850*** | -19.226*** | .067 |

Notes: All variables are in log. ADF and PP are the augmented Dickey–Fuller (1979) test and Phillips–Perron (1988) test, respectively. KPSS is Kwiatkowski–Phillips–Schmidt–Shin (1992) test. Lag length for ADF was chosen by Schwarz information criterion (SIC). Bandwidth for PP and KPSS using Bartlett Kernel.

**Statistical significance levels for the rejection of the null hypothesis at the 5%.

***Statistical significance levels for the rejection of the null hypothesis at the 1%.

Table 3. Johansen cointegration test on revenues and spending without break

| Number of cointegrating equations | Null | Statistic | Critical value |
|---|------------|-----------|----------------|
| Trace statistic (λ_{trace}) | $r = 0$ | 13.383 | 15.495 |
| | $r \leq 1$ | .0008 | 3.845 |
| Maximum eigenvalue statistic (λ_{max}) | $r = 0$ | 13.382 | 14.265 |
| | $r \leq 1$ | .0008 | 3.845 |

Notes: Lag length for Johansen cointegration tests are chosen based on Schwarz Information criterion (SIC).

where $DU1_t$ and $DU2_t$ are dummy indicator variables for a mean shift occurring at times $TB1_t$ and $TB2_t$, and $DT1_t$ and $DT2_t$. All the remaining variables are as defined in Equation (6). The results of the unit root tests are presented in Tables 1–3. The critical values of this test are tabulated in LP (1997).

3.2. Standard cointegration tests

The decision on the appropriate cointegration method to be used for examining the long run relationship between revenues and expenditures will depend on whether there is a structural break(s) in the unit root tests' results or not. If, for example, breaks are not found, and the series are integrated of order one $I(1)$, the analysis will employ the Johansen (1995) cointegration technique.¹ This cointegration approach is based on the vector autoregressive (VAR) models. Considering a VAR of order:

$$y_t = A_1 y_{t-1} + \dots + A_p y_{t-p} + Bx_t + \varepsilon_t \quad (8)$$

where y_t is a k -vector of nonstationary $I(1)$ variables, x_t is a d -vector of deterministic variables, A_1, \dots, A_p and B are matrices of coefficients to be estimated, and ε_t is a vector of innovations, the VAR equation can be rewritten as

$$\Delta y_t = \Pi y_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta y_{t-i} + Bx_t + \varepsilon_t \quad (9)$$

where, $\Pi = \sum_{i=1}^p A_i - 1$, $\Gamma = -\sum_{j=i+1}^p A_j y_t$ is an $m \times 1$ vector of endogenous, $I(1)$ variables (R and G), Π , Γ_i and B are parameters to be estimated, and x_t is a $q \times 1$ vector of the exogenous or deterministic $I(0)$ elements (constant and trend), and ε_t is the matrix of usual random errors that follows Gaussian white noise with zero mean and constant variances.

The Johansen test centers on an examination of the Π matrix, which contains both long run information and the speed of adjustment. The test for cointegration between the y_s is calculated by looking at the rank of the Π matrix via its eigenvalue. For the Johansen procedure, there are two test statistics for the number of cointegrating vectors: the trace and the maximum eigenvalue value statistics. In the present study, both the trace (λ_{trace}) tests and the maximum eigenvalue (λ_{max}) statistics are used to estimate the cointegration rank r (the number of independent cointegrating vectors):

$$\lambda_{\text{trace}}(r) = -T \sum_{i=r+1}^n \ln(1 - \hat{\lambda}_i^2)$$

$$\lambda_{\text{max}}(r, r+1) = T \ln(1 - \lambda_{r+1})$$

where λ_i denotes the eigenvalue obtained from the estimated matrix and T signifies the number of observations to be used after lag adjustment. The trace test (λ_{trace}) is based on the null hypothesis that the number of cointegrating vectors is less than or equal to r against the general alternative. The maximum eigenvalue test (λ_{max}) is similar, except that the alternative hypothesis is explicit. The null hypothesis r is tested against the alternative $r+1$.

Although the standard cointegration tests assumed that the cointegrating vectors are time invariant, it is argued that the Johansen (1995) test may lose power when faced with one structural break

or more (see, e.g. Johansen et al., 2000). Thus, accounting for a possible break or breaks in the observed data is pertinent, especially to correct for the potential shift in their cointegrating vector(s).

3.3. Cointegration test with structural breaks

Johansen et al. (2000) developed an alternative cointegration test that extends the standard Johansen and Juselius (1990) framework to allow for possible exogenous breaks in trends and levels of the deterministic components of a vector-valued stochastic process. To test cointegration relations among p time series, Johansen et al. (2000) divide a sample of T observation points into q subsamples according to the position of the exogenous break points denoted by $v_j = (T_j/T)$, where T is the full sample size and T_j represents the last observation of the j th subsample; $j = 0, 1, 2, \dots, q$, with $T_0 = 1$ and $T_q = T$. This technique disregards observations after the possible q exogenous breaks by including a number of impulse dummies and then use two variants of the standard trace test, called the $H_l(r)$ and $H_c(r)$ tests for which there are $(q - 1)$ breaks, where r represents the cointegrating rank (see also Giles & Godwin, 2012; Li & Daly, 2009 among others). This method maintained a level VAR model of order k , and is formulated conditionally on the first k observation of each subsample, $y_{T_{j-1}+1}, \dots, y_{T_{j-1}+k}$ and for $j = 1, 2, \dots, q$, the observation $T_{j-1} + k \leq t \leq T_j$ is described by the following equation

$$\Delta y_t = (\Pi, \Pi_j) \begin{pmatrix} y_{t-1} \\ t \end{pmatrix} + \mu_j + \sum_{i=1}^{k-i} \Gamma_i \Delta y_{t-i} + \varepsilon_t \quad (10)$$

$j = 1, 2, \dots, q$, where Π_j and μ_j are $(p \times 1)$ vectors. ε_t are $(p \times 1)$ vectors of disturbance terms assumed to be purely random. As in the standard Johansen cointegration approach, the cointegration is signified by the restriction $\Pi = \alpha\beta$, where α and β are $(p \times r)$ full rank matrices. To rule out the possibility of quadratic trends in the p time series, the assumption that $\Pi_j = \alpha\gamma_j$ is maintained in each subsample, where γ_j is $(l \times r)$ full rank matrix. The intervention dummy variables for the breakpoints $(q - 1)$ can be specified as (see also Giles & Godwin, 2012; Joyeux, 2007; Li & Daly, 2009)

$$D_{j,t} = \begin{cases} 1 & \text{for } T_{j-1} + 1 \leq t \leq T_j, \\ 0 & \text{otherwise} \end{cases} \quad j = 1, 2, \dots, q.$$

and

$$E_{j,t} = \sum_{i=k+1}^{T_j-T_{j-1}} D_{j,t-i} \begin{cases} 1 & \text{for } t = T_{j-1} + 1 \leq t \leq T_j, \\ 0 & \text{otherwise} \end{cases} \quad j = 1, 2, \dots, q.$$

Every $D_{j,t}$ represents an indicator for the end of $(j - 1)$ th subsample, while its lagged values indicate the individual observation points within the j th subsample. Summing up these lagged values makes $E_{j,t}$ an indicator dummy for the entire subsample. Using these intervention dummies, the individual models of the whole subsamples can be combined into a single equation as follows (see Joyeux, 2007; Li & Daly, 2009)

$$\Delta y_t = \alpha \begin{bmatrix} \beta \\ \gamma \end{bmatrix}' \begin{bmatrix} y_{t-1} \\ tE_t \end{bmatrix} + \mu E_t + \sum_{i=1}^{k-1} \Gamma_i \Delta y_{t-i} + \sum_{i=1}^k \sum_{j=2}^q k_{ji} D_{j,t-i} + \varepsilon_t \quad (11)$$

where μ is the $(p \times q)$ matrix, γ is $(q \times r)$ and E_t is $(q \times 1)$ matrix, and k_{ji} is $(p \times 1)$ vectors. Johansen et al. (2000) proposed a maximum likelihood cointegration test method based on the squared sample canonical correlations, $\hat{\lambda}_i$ of Δy_t and $(\hat{y}_{t-1}, tE_t)'$ corrected for the regressors

$$E_t, \quad \Delta y_{t-i}, \quad i = 1, \dots, k - 1, \quad D_{j,t-i}, \quad i = 1, \dots, k; j = 2, \dots, q.$$

The likelihood ratio test statistic for this model is specified as (Giles & Godwin, 2012)

$$LR\{H_l(r)/H_l(p)\} = -T \sum_{i=r+1}^p \log(1 - \hat{\lambda}_{li})$$

and

$$LR\{H_c(r)/H_c(p)\} = -T \sum_{i=r+1}^p \log(1 - \hat{\lambda}_{ci})$$

where $\hat{\lambda}_{li}$ signify the squared canonical correlations, corrected for the regressors in the broken linear trend model, while the values of $\hat{\lambda}_{ci}$ are the corresponding canonical correlations, corrected for the regressors in the broken constant level model.²

The moments (mean and variance) of the asymptotic distribution can be obtained by the estimated response surfaces presented in Table (4) of Johansen et al. (2000). These moments can then be used to select suitable parameters for a gamma distribution. From these, the critical values for the test statistics ($H_l(r)$ and $H_c(r)$) can be calculated, which in turn depend on the number and location of the break points. It is essential to remark that the critical values also depend on $(p - r)$, where p represents variables (revenues and expenditures, $p = 2$, here) under test, and r denotes the cointegrating rank being tested. Consequently, for the present analysis, $r = 0, 1$.

It is worth noting that since the application of Johansen et al. (2000) requires that the breakpoint(s) is/are known a posteriori, the analysis employs the technique proposed by Bai and Perron (1998, 2003) to identify the possible breakpoints in a long run relationship between the observable series.³

3.4. Multiple breakpoint tests

The Bai and Perron technique extended the basic Chow test and developed a model with m breaks ($m + 1$ regimes) to detect unknown multiple structural breaks. The model is formally presented as

$$y_t = x_t' \beta + \zeta_t' \delta_j + \mu_t, \quad t = T_{j-1} + 1, \dots, T_j \quad (12)$$

for $j = 1, \dots, m + 1$. y_t denotes the dependent variable observed at time t . x_t' and ζ_t' are $(p \times 1)$ and $(q \times 1)$ vectors of the explanatory variables; β and δ_j are vectors of coefficients associated with x_t' and ζ_t' , respectively. μ_t is the usual error term assumed to be independently and identically distributed with zero mean and constant variance. The break dates, denoted by (T_1, \dots, T_m) , are explicitly unknown a priori, with $T_0 = 0$ and $T_{m+1} = T$.

Table 4. Zivot and Andrews test for unit roots with one structural break

| | R_t | ΔR_t | G_t | ΔG_t | BD_t | ΔBD_t |
|-----------------|-----------------|--------------------|------------------|--------------------|--------------------|--------------------|
| TB | 1996 | 1996 | 1996 | 1996 | 2000 | 2000 |
| α | -.5922 (-4.110) | -1.514*** (-7.861) | -1.291 (-4.010) | -4.239*** (-6.550) | -.8640*** (-7.045) | -2.357*** (-7.603) |
| θ | .1470* (4.958) | .0826 (3.207) | .2451*** (6.231) | .1508 (4.790) | .0585* (4.836) | .0454 (3.119) |
| γ | .0040 (1.321) | -.0042 (-2.099) | .0075 (1.453) | -.0087 (-3.667) | -.0004 (-.3864) | -.0033 (-2.377) |
| K | 2 | 1 | 7 | 6 | 0 | 0 |
| Critical values | | | | | | |
| 1% | -5.57 | -5.57 | -5.57 | -5.57 | -5.57 | -5.57 |
| 5% | -5.08 | -5.08 | -5.08 | -5.08 | -5.08 | -5.08 |
| 10% | -4.82 | -4.82 | -4.82 | -4.82 | -4.82 | -4.82 |

Note: Lag length for the test was chosen by Schwarz information criterion (SIC). Figures in brackets are t-statistics.

*Statistical significance at the 10% level.

**Statistical significance at the 5% level.

***Statistical significance at the 1% level.

Equation (12) can be re-written in a matrix form as

$$Y = X\beta + \bar{Z}\delta + U, \quad (12.1)$$

where $Y = (y_1, \dots, y_T)'$, $X = (x_1, \dots, x_T)'$, $U = (u_1, \dots, u_T)'$, $\delta = (\delta'_1, \delta'_2, \dots, \delta'_{m+1})'$, and \bar{Z} is the matrix which partitions Z diagonally at (T_1, \dots, T_m) , that is $Z_t = (Z, \dots, Z_{m+1})$ with $Z_t = (\zeta_{T_{i-1}+1}, \dots, \zeta_{T_i})'$. To obtain the estimates of the coefficients in (12) and determine all possible sets of m breakpoints, the method uses a global optimization procedure to minimize the sum of square residuals of the least square model (12) as follows (Bai & Perron, 1998, 2003)

$$(Y - X\beta + \bar{Z}\delta)'(Y - X\beta + \bar{Z}\delta) = \sum_{i=1}^M \sum_{t=T_{i-1}+1}^{T_i} [y_t - x'_t\beta - \zeta'_t\delta_i]^2. \quad (12.2)$$

Estimates from this procedure are then used as a basis for some proposed breakpoint tests that include the following (Bai & Perron, 1998, 2003).

The supF type tests the null hypothesis of no structural break, $m = 0$ vs. $m = k$ breaks. For every partition, (T_1, \dots, T_k) , the null hypothesis of no breaks is tested on the full sample observation, while the alternative hypothesis is tested against an (pre-specified) alternative of breaks that involves estimation of every subsample of dimension $T_1 = [T\lambda_i]$ ($i = 1, \dots, k$). To evaluate the null hypothesis that $\delta_1 = \delta_2 = \dots = \delta_{k+1}$, Bai and Perron (2003) proposed the following F -statistic

$$F_T(\lambda_1, \dots, \lambda_k; q) = \frac{1}{T} \left(\frac{T - (k+1)q - p}{kp} \right) \hat{\delta}' R(\hat{V}(\hat{\delta})R')^{-1} R\hat{\delta} \quad (12.3)$$

where $\hat{\delta}$ is the optimal k -break estimate of δ , $(R\delta)' = (\delta'_1 - \delta'_2, \dots, \delta'_k - \delta'_{k+1})$, and $\hat{V}(\hat{\delta})$ is an estimate of the variance covariance matrix of $\hat{\delta}$ that may be robust to serial correlation and heteroskedasticity.⁴

To test the null hypothesis of no structural break against an alternative of unknown k breakpoints up to some upper bound, m , Bai and Perron suggested the following specifications called the double maximum tests:

The first test which is termed the equal-weighted version takes the form $UD \max F_T(Mq) = \max_{1 \leq m \leq M} F_T(\hat{\lambda}_1, \dots, \hat{\lambda}_m; q)$, where $\hat{\lambda}_j = \hat{T}_j/T$ ($j = 1, \dots, m$); the estimates of the breakpoints are obtained using the global optimizer specified above. The second test which applies weights to the individual tests such that the marginal probability values are identical across m values has the form $WD \max F_T(Mq)$; details are provided in Bai and Perron (1998).

Bai and Perron (1998) also proposed an alternative test termed $\sup F_T(l+1/l)$. The method tests the null hypothesis l breaks against the alternative hypothesis of $l+1$ breaks to each subsample containing the observations \hat{T}_{i-1} to \hat{T}_i ($i = 1, \dots, l+1$). For details on the critical values for these tests, see Bai and Perron (1998, 2003).

Bai and Perron, however, note that upon all the three techniques, the sequential method works better. Thus, the present analysis relies upon the sequential procedure to interpret the structural break(s) test results.

Although the system-based approach to cointegration such as the Johansen vector error correction (VECM) maximum-likelihood estimator is particularly suitable when there are multiple cointegrating relationships between the observable series, the method is extremely sensitive to the choice of lag length (see Chang & Caudill, 2005; Jahangir Alam, Ara Begum, Buysse, & Van Huylbroeck, 2012; Muscatelli & Spinelli, 2000). Also, Stock and Watson (1993) show that system-based estimators tend to exhibit greater dispersion compared to the single-equation estimators. It has also been

demonstrated that unlike the multivariate frameworks, univariate estimators such as an autoregressive distributed lag (ARDL) and the dynamic ordinary least squares (DOLS) models tend to have minimum biases in a relatively small sample (see Narayan & Narayan, 2005; Odhiambo, 2009; Pesaran & Shin, 1999; Stock & Watson, 1993).

Consequently, as a robustness check, and to formally assess the strength of the budget deficit sustainability, the analysis will further re-estimate the long run relationship between government revenues and spending using the bounds testing approach to cointegration.⁵ This method, developed by Pesaran and Pesaran (1997) and Pesaran, Shin, and Smith (2001), can be implemented in the form the conditional ARDL model. One obvious advantage of this method is that it does not require pre-testing the integration order of the variables. This eliminates the problem of low power associated with conventional unit root tests such as, for example, ADF and PP tests (Wang & Tomek, 2007). Thus, the approach is applicable regardless of whether the regressors in Equation (6) are purely $I(0)$, purely $I(1)$, or are mutually cointegrated. In what follows, Equation (6) is estimated based on the following error correction regression (Pesaran & Pesaran, 1997; Pesaran et al., 2001)

$$\Delta R_t = \delta_0 + \psi_R R_{t-1} + \psi_G G_{t-1} + \sum_{i=1}^p \varpi_i \Delta R_{t-i} + \sum_{i=0}^p \lambda_i \Delta G_{t-i} + \varepsilon_{1t} \quad (13)$$

where Δ denotes the first difference operator. All the remaining variables are as defined in Equation (5). The lag orders in the ARDL model are chosen based on Schwarz information criterion (SIC), and the selected model can be estimated using an ordinary least squares technique (OLS). The bounds test for examining the presence of a long run relationship among the variables of interest can be conducted using the F -test statistic, which tests the joint significance of the coefficients of the lagged-level variables. So, the null hypothesis in Equation (13) is ($H_0: \psi_R = \psi_G = 0$) denoted by $R_R(R/G)$ against the alternative hypothesis ($H_1: \psi_R \neq \psi_G \neq 0$).

The F -test that can be used to examine these hypotheses has a nonstandard distribution with two sets of critical values for certain significance levels in Pesaran et al. (2001). The first set of critical values assumes that all the variables are $I(0)$ and the second set assumes that all variables are $I(1)$. These critical values depend on the sample size; the number of explanatory variables; whether the series are $I(0)$ or $I(1)$; and whether the ARDL model contains an intercept and/or a trend (see, e.g. Narayan, 2005). If the estimated F -statistics falls outside the upper bounds of $I(1)$, then a conclusion can be made that cointegration exists among the variables. If the computed F -statistics falls below the lower bound of $I(0)$, then the null of no cointegration cannot be rejected. If the calculated F -statistics falls in between the two bounds, then the test becomes inconclusive.

Further robustness of the long run relationship between the government revenues and spending will also be re-examined using the DOLS technique.

The DOLS estimator proposed by Stock and Watson (1993) extends the traditional (static) OLS regression by employing lags, leads, and contemporaneous values of the explanatory variable in first difference. One significant advantage of using DOLS is that it produces asymptotically efficient estimates for cointegrated variables even in a small sample (Narayan & Narayan, 2005; Stock & Watson, 1993). Let the OLS model take the form

$$y_t = \alpha_0 + x_t' \beta + \varepsilon_t \quad (14)$$

where β is a $(K \times 1)$ vector of the slopes of the regressors, x_t' is a $(K \times 1)$ vector of the autoregressive process of the first-order difference of the explanatory variables: $x_t' = x_{t-1}' + \varepsilon_t$, and ε_t is the usual error term. In what follows, the DOLS estimator of (4) can be expressed as

$$y_t = \alpha_0 + x_t' \beta + \sum_{i=-l_1}^{i=l_2} \varphi_i \Delta x_{t+i} + v_t \quad (15)$$

where φ_i is the coefficient of a lag and lead of the first differenced regressors. Suppose y_t is found to be $I(1)$ and at least one of the explanatory variables is $I(1)$ or $I(0)$; the DOLS estimates are obtained by regression of Equation (15).

In the presence of cointegration between government revenues and spending, it is essential to include an error correction term (ECT) in the stationary model to capture the short run deviations of the series from their long run equilibrium (see Engle & Granger, 1987). Accordingly, the present analysis is extended to test the alternative hypothesis regarding the revenues and spending relationships in Nigeria as follows

$$\Delta R_t = \delta_{10} + \sum_{i=1}^p \varpi_{1i} \Delta R_{t-i} + \sum_{i=1}^p \lambda_{1i} \Delta G_{t-i} + \eta_1 \hat{\varepsilon}_{t-1} + v_{1t} \quad (16)$$

$$\Delta G_t = \delta_{20} + \sum_{i=1}^p \varpi_{2i} \Delta G_{t-i} + \sum_{i=1}^p \lambda_{2i} \Delta R_{t-i} + \eta_2 \hat{\varepsilon}_{t-1} + v_{2t} \quad (17)$$

where $\hat{\varepsilon}_{t-1}$ is the lagged ECT from the cointegrating relationship between the government revenues and spending. $\varpi_1, \varpi_2, \lambda_1, \lambda_2, \eta_1$ and η_2 are the parameters to be estimated. η_1 and η_2 are the adjustment coefficients that capture speed at which any short run deviations of the observable series revert to the long run equilibrium. The standard assumption in econometric modeling requires an adjustment coefficient to be negative and statistically significant. The short run causality can be determined by the statistical significance of the partial F -statistic in each of Equations (16) and (17), while the long run causality can be determined by the statistical significance of the corresponding ECT in each equation.

It should be noted, however, that the fundamental assumption of the standard cointegration techniques discussed above is that the adjustment behavior of government revenues and spending due to a budgetary disequilibrium is that the relationship between the series is linear. Enders and Siklos (2001), however, note that such assumption may be misleading as it will likely result in model misspecification if the actual relations are asymmetric.

Notwithstanding, the robustness of the findings presented above, the assumption of a linear relationship between the two variables may lead to erroneous fiscal policy prescriptions and adverse consequences if the actual adjustment of the variables is asymmetric over budgetary disequilibria (deficits and surpluses). In fact, studies have documented that key macroeconomic variables tend to exhibit asymmetric adjustment through business cycles, especially when the closeness between the budgetary process and the business cycle is taken into consideration (see, e.g. Ewing, Payne, Thompson, & Al-Zoubi, 2006; Jibrilla & Mohammed, 2015).

The subsequent analysis will, therefore, require a more general model that allows nonlinearity in the cointegration relationship between revenues and spending. In what follows, the techniques proposed by Enders and Siklos (2001) are used to examine the possible nonlinear relationship between the series. This procedure, which is a modified version of Eagle and Granger two-step cointegration technique (1987) in the form of threshold autoregression (TAR) model and momentum threshold autoregression (M-TAR) model, is based on the Tong (1990).⁶

Based on these techniques, the analysis partitions lagged sequence of residuals generated from the long run relationship between revenues and spending as follows:

$$\Delta \hat{\varepsilon}_t = I_t \rho_1 \hat{\varepsilon}_{t-1} + (1 - I_t) \rho_2 \hat{\varepsilon}_{t-1} + \sum_{i=1}^{p-1} \lambda_i \hat{\varepsilon}_{t-i} + \mu_t \quad (18)$$

where $\mu_t \sim \text{I.I.D}(0, \sigma^2)$ and the lagged values of $\Delta\hat{\varepsilon}_t$ are expected to produce uncorrelated residuals. I_t is the Heaviside indicator function associated with the TAR and M-TAR, which are specified as (19) and (20), respectively.

$$I_t = \begin{cases} 1 & \text{if } \hat{\varepsilon}_{t-1} \geq \tau \\ 0 & \text{if } \hat{\varepsilon}_{t-1} < \tau \end{cases} \quad (19)$$

$$I_t = \begin{cases} 1 & \text{if } \Delta\hat{\varepsilon}_{t-1} \geq \tau \\ 0 & \text{if } \Delta\hat{\varepsilon}_{t-1} < \tau \end{cases} \quad (20)$$

The stationarity of the sequence, $\hat{\varepsilon}$ (or $\Delta\hat{\varepsilon}_t$), requires $-2 < (\rho_1, \rho_2) < 0$, and the threshold value is endogenously determined using Chan's (1993) method. This method arranges the values of the TAR and M-TAR models in an ascending order and eliminates the largest and smallest 15%. If the deviation of (or $\Delta\hat{\varepsilon}_t$) is above the threshold, the adjustment is represented by $\rho_1\hat{\varepsilon}_{t-1}$ (or $\rho_1\Delta\hat{\varepsilon}_t$), while the adjustment for the deviation of $\hat{\varepsilon}_{t-1}$ below threshold is denoted by $\rho_2\hat{\varepsilon}_{t-1}$ (or $\rho_2\Delta\hat{\varepsilon}_t$). These adjustments are represented by dummy values: the indicator will take the value of 1 for deviation above threshold and 0 for deviation below a threshold value. Whether positive and negative divergences have different effects on the behavior of government revenues–spending nexus could be determined by the estimated values of ρ_1 and ρ_2 . For instance, if $|\rho_1| < |\rho_2|$, the adjustment is slow for deviation above threshold value.

In the case of an adjustment that unveils more persistence whenever the sequence $\hat{\varepsilon}_{t-1}$ (or $\Delta\hat{\varepsilon}_{t-1}$) < 0 in a TAR (or M-TAR) model Chan (1993) showed that a super-consistent estimate of the threshold can be found by searching over all values of the lagged residuals sequence. This is to minimize the sum of squares errors (SSE) from the fitted threshold model(s). The present analysis follows the standard procedure of using only 70% of the sample observations as potential thresholds.

The null hypothesis of no cointegration for both TAR and M-TAR models can be denoted as $\rho_1 = \rho_2 = 0$. The F -statistics to examine this null has a nonstandard distribution with certain critical values (Tables 1 and 2) in Enders and Siklos (2001). Rejection of the null hypothesis implies that either ρ_1 or ρ_2 is at least significantly greater than zero. This then allows the test for the presence of linear adjustment processes. This can be done by setting the null hypothesis as $\rho_1 = \rho_2$, which can be tested using the standard F -test (or Fisher test) statistic. However, if this null is rejected, one can conclude that the cointegration relationship between the two variables is nonlinear and the adjustment is asymmetric.

3.5. Data

The Nigerian data relating to the time series of the real government revenues, real spending, and the real budget deficit (BD) were taken from the Central Bank of Nigeria Statistical Bulletin (various issues) over 1961–2014 period. All data are scaled as a share of real GDP. The study is aimed to analyze how the relationship between government revenues and spending in the last 54 years influenced budget deficit sustainability in Nigeria. The start and end periods of our sample are dictated by data availability, as no data are thus far obtainable for the year 2015 (at the time of writing this report). Plots of log trends and logs of the variables used are presented in Figures 2 and 3, respectively.

4. Empirical results

4.1. The standard unit root and cointegration tests' results

To allow for more general specifications, the standard (ADF, PP, and KPSS) unit root tests' equations include both constant terms and trends in the levels of the variables, while only constant is included in their first difference. The test results are shown in Table 2.

Figure 2. Trends of public revenues and spending (in billion naira), 1961–2014.

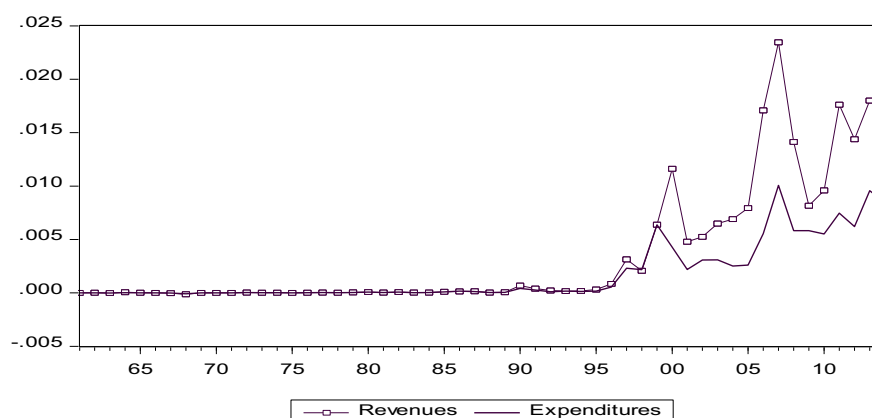
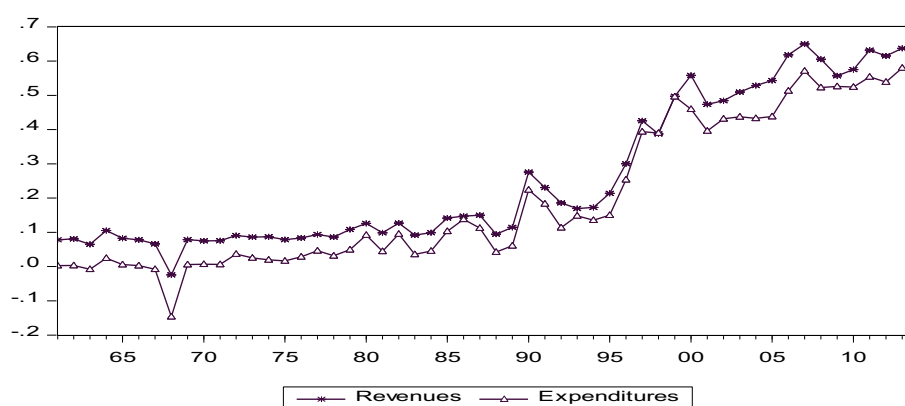


Figure 3. Logarithmic of public revenues and spending, 1961–2014.



As reported in Table 2, both R_t and G_t are found to be nonstationary at their levels across all the alternative tests, but are stationary at their first differences, hence are integrated of order one, $I(1)$ and are robust across all the alternative unit root tests used. Whereas for BD_t , all test statistics from the three alternative unit root tests show that the budget deficit variable is integrated of order zero, $I(0)$. The absence of a unit root in BD_t tends to suggest that the budget deficit may be sustainable, at least as far as can be detected in the sample used.

Next, the analysis proceeds to verify the sustainability of the Nigerian budget deficit by examining the long run relationship between revenues and spending using Johansen (1995) multivariate cointegration technique. The purpose of the cointegration test is to determine whether the $I(1)$ variables are cointegrated, which according to Johansen–Juselius (1990) technique require their linear combination to produce $I(0)$. The Johansen–Juselius cointegration approach assumes no trend in the cointegrating equation (Johansen, 1995). Results are reported in Table 3.

As can be observed from the estimated results of the Johansen cointegration test (Table 3), both the trace test and the maximum eigenvalue tests indicate no cointegrating equation at the conventional significance level, suggesting the absence of a long run relationship between government revenues and spending in Nigeria over the sample period. The results are not too surprising as the lack of cointegration might be due to the presence of structural break(s) in their long run relationship. To account for the possible break(s) in the data-set, the analysis begins by re-examining their integration order using the methodology proposed by ZA (1992), as discussed in Section 3.1. The unit root tests' results for ZA test with one endogenous structural break are reported in Table 4.

4.2. The results of the unit root test with breaks

As evident in Table 4, when a possible break in the data-set is accounted for, the ZA unit root test results suggest that only budget deficit exhibits stationarity properties.⁷ However, the null hypothesis of a unit root is rejected for both government revenues and spending at their first differences, signifying that they are integrated of order one, $I(1)$. This result is consistent with the standard ADF, PP, and KPSS tests' results.

Though the ZA (1992) test is an improvement over standard ADF, PP, and KPSS tests, it is contended that the ZA test may lose power when confronted with more than a single break (see Narayan, 2005). In an attempt to tackle this issue, as previously explained, LP (1997) extends ZA methodology and developed a technique that accounts for two endogenous breaks. The LP test results are reported in Table 5.

The LP unit root test with two endogenous structural breaks provides results which are similar to the ZA tests: only reject the null hypothesis of a unit root for both government revenues and spending at their first difference, while the budget deficit is found to be stationary (albeit weak at level).

Since both R_t and G_t are integrated of order one, $I(1)$ with breaks, the analysis proceeds to identify further the possible breakpoints in their long run relationship. Table 6 presents the test results.

As shown in Table 6, the Bai and Perron (2003) multiple break test results reject the null hypothesis of no break(s) in the long run relationship between revenues and spending in Nigeria with breaks in 1990 and 2000 at 1% significance level.

The first breakpoint which occurred in 1990 appears to coincide with the resultant effect of the Gulf War (Iraqi's invasion of Kuwait) on the international oil market. The war pushes the oil price in 1990 to rise by about 33% compared to 1989. Nigeria used the revenue windfalls from this increase to expand public sector expenditure. Besides, due to the country's over dependence on oil revenues,

Table 5. Lumsdaine and Papell test for unit roots with two breaks

| | R_t | ΔR_t | G_t | ΔG_t | BD_t | ΔBD_t |
|-----------------|-----------------|--------------------|-----------------|-------------------|------------------|--------------------|
| TB ₁ | 1968 | 1996 | 1968 | 1996 | 1971 | 2000 |
| TB ₂ | 1997 | 2001 | 1997 | 2000 | 2000 | 2007 |
| α | -.9189 (-5.902) | -1.622*** (-9.302) | -1.325 (-5.838) | -2.271** (-6.884) | -1.021* (-6.658) | -2.516*** (-8.623) |
| θ | -.0461 (-1.997) | .1205 (3.043) | -.0562 (-1.672) | .0777 (1.532) | -.0167* (-1.753) | .0581 (3.343) |
| ω | .1923 (5.370) | -.0938 (-2.648) | .2369 (6.191) | -.1644 (-3.779) | .0571 (4.633) | -.0479 (-2.259) |
| γ | .0059 (5.019) | -.0051 (-4.460) | .0097 (5.651) | .0229 (1.148) | -.0012 (-2.919) | -.0057 (-1.505) |
| ψ | .0060 (2.309) | .0062 (5.250) | .0072 (2.093) | -.0226 (-1.1078) | -.0001 (-.0958) | .0112 (2.317) |
| K | 1 | 1 | 4 | 3 | 1 | 2 |
| Critical values | | | | | | |
| 1% | -7.34 | -7.34 | -7.34 | -7.34 | -7.34 | -7.34 |
| 5% | -6.82 | -6.82 | -6.82 | -6.82 | -6.82 | -6.82 |
| 10% | -6.49 | -6.49 | -6.49 | -6.49 | -6.49 | -6.49 |

Note: Lag length for the test was chosen by Schwarz information criterion (SIC). Figures in brackets are t -statistics.

*Statistical significance at the 10% level.

**Statistical significance at the 5% level.

***Statistical significance at the 1% level.

the resulting vulnerability to oil price shock has since the pre-1990 (particularly since the early 1980s) increased her expenditure volatility (see Budina et al., 2007). In fact, the World Bank (1997) notes that notwithstanding the sharp rise in the World oil price (though was much shorter-lived), the country's fiscal deficit rose in 1990. Available data indicate that fiscal deficits and national debt have increased by about 161 and 30% from 1989, respectively, apparently making it consistent with the breakpoint.⁸

The federal budget had also witnessed some significant changes between 2000 and 2001. A substantial increase in government spending occurred in these years due to a sharp rise in crude oil price in 2000. However, such an expansion in spending immediately leads to macroeconomic instability as inflation rate increased from less than 7% in 1999 to more than 18% in 2001 and exchange rate became very unstable (see, e.g. Jibrilla & Muhammad Dodo, 2014; OECD/AfDB, 2002). These circumstances might be responsible for the break obtained i The bounds test results are presented n 2000.

Next, the study employs the methods introduced by Johansen et al. (2000). Since the sequential break tests (Table 6) provide evidence that Nigerian "budgetary process" experienced significant breaks in 1990 and 2000, linked to fiscal resource volatility, as already discussed, following Johansen et al. (2000), $q = 3$ with $v_1 = T_1/T = .52$, and $v_2 = T_2/T = .70$. The test results are reported in Table 7.

Table 7 shows that when the deterministic components of revenues and spending allowed structural breaks at 1990 and 2000, the Johansen cointegration test results reject the null hypothesis of no cointegration at better than 5% significance level. This result suggests sustainability of the budget deficit in Nigeria in the long run. As mentioned in Section 3, the robustness of Johansen et al. (2000) cointegration is checked using ARDL (bounds testing) cointegration technique.⁹ The bounds test results are presented in Table 8.

Table 6. Bai and Perron breakpoints test

| h = 15 | Test statistics | Critical value** |
|---------------------------------------|---------------------------|------------------|
| | Sup $F_T(1)$ 23.037** | 11.47 |
| | Sup $F(2/1)$ 33.380** | 12.95 |
| | Sup $F_T(2)$ 50.760** | 9.75 |
| | UDmax101.519** | 11.70 |
| | WDmax119.428** | 12.81 |
| Coefficient estimates with two breaks | | |
| | $\hat{\delta}_1$ | $\hat{\delta}_2$ |
| | .8599 (.0363) | .9016 (.0576) |
| | Number of breaks selected | |
| Sequential | 2 | |
| LWZ | 2 | |
| SIC | 2 | |
| Corresponding estimated break dates | | |
| | \hat{t}_1 | \hat{t}_2 |
| | 1990 [1990–1999] | 2000 [2000–2014] |

Notes: Test statistics employ HAC covariances (Prewitening with lags = 1, Quadratic-Spectral kernel, Andrews bandwidth). Estimation allows heterogeneous error distributions across breaks. Figures in brackets under the coefficient estimates are standard errors.

**Bai–Perron (2003) critical values and statistical significance at the 5% level.

Table 7. Johansen cointegration test on revenues and spending with breaks in 1990 and 2000

| Number of cointegrating equations | Null | Statistic | 10% critical value | 5% critical value | 1% critical Value |
|-----------------------------------|------------|-----------|--------------------|-------------------|-------------------|
| $H_l(r)$ | $r = 0$ | 54.258*** | 42.999 | 46.355 | 53.098 |
| | $r \leq 1$ | 20.239 | 21.468 | 24.027 | 29.335 |
| $H_c(r)$ | $r = 0$ | 34.019 | 29.046 | 31.828 | 37.496 |
| | $r \leq 1$ | 20.239** | 13.918 | 15.970 | 20.312 |

Note: Lag length for the test was chosen by Schwarz information criterion (SIC).

**Statistical significance at the 5% level.

***Statistical significance at the 1% level.

Table 8. ADRL bounds test results for cointegration relationship

| Critical value of the F -statistic for the bounds test results with intercept and no time trend | | | | | | |
|---|-------|-------|-------|-------|-------|-------|
| k | 1% | | 5% | | 10% | |
| | I(0) | I(1) | I(0) | I(1) | I(0) | I(1) |
| 3 | 5.333 | 7.063 | 3.710 | 5.018 | 3.008 | 4.150 |
| Computed F -statistics | | | | | | |
| $F_R(R/G)$ | 5.538 | | | | | |

Note: Critical values are from Narayan (2005).

The calculated F -statistic $R_R(R/G) = 5.538$ for model (9) appears to be higher than the critical value (5.018) of the upper bound at the 5% significance level (Table 8). See also Appendix A for a cointegration graph that shows the linear combination of the government revenues and spending to be $I(0)$. Thus, rejecting the null hypothesis of no cointegration between government revenues and spending confirms the results reported in Table 7. However, as noted earlier, the strength of the budget deficit sustainability requires not only cointegration between revenues and spending, but also the magnitude of the slope or coefficient relating the two series.

Therefore, having verified a cointegration relation between R_t and G_t , the analysis proceeds to examine the long run coefficients using the ARDL methodology.

The results from the long run models estimated using the ARDL technique, which were also re-estimated using DOLS as a robustness check, are presented in Table 9.¹⁰

From these results, it can be noted that the slope parameter (coefficient on the government spending) is significantly less than one, suggesting that the stock of public debt, on average, appears to grow faster than the growth rate of the real GDP. This finding tended to indicate that the government may have been following a SBC strategy (see Section 1 for the likely cause(s) of this finding). Also, the coefficient on the dummy variables is statistically different from zero (except D1 in DOLS Equation), confirming the existence of structural breaks.

To ascertain the direction of causality, the Granger casualty test has been carried out and the results are presented in Table 10.

To find out whether the estimated long run models (presented in Table 9) were correctly specified or not, diagnostics of normality, serial correlation, and heteroskedasticity tests were conducted. As can be observed in Panel B of Table 10, the test results are in conformity with the classical assumptions. Besides, the stability tests of recursive residuals using CUSUM and CUSUM Squares (Appendix B) indicate that the estimated ARDL model is free from misspecification, and all the coefficient

estimates are relatively stable, respectively, at the 5% significance level. Moreover, Hansen instability test (Panel B of Table 9) provides evidence for parameter stability of the long run (DOLS) estimates at conventional levels of significance.

The Granger causality results between government revenues and spending are presented in Panel A of Table 10.¹¹ In the short run, the results indicate each of the variables has a positive and statistically significant effect on one another. Also, the partial *F*-statistics in both the revenues and spending equations indicate short run bidirectional Granger causality between the government revenues and spending in Nigeria.

The evidence, which shows bidirectional causality between government revenues and expenditures both in the short run and long run, provides support for the fiscal synchronization hypothesis in Nigeria. Besides, the findings tended to be consistent with that of Kollias and Makrydakis (2000) for Greece and Ireland, Li (2001) for China, Nyamongo et al. (2007) for South Africa, and Aregbeyen and Ibrahim (2012) for Nigeria, among others, and provided evidence against Obioma and Ozughalu (2010) for the case of Nigeria. A likely justification is that this study does not only allow for structural breaks in the cointegration and causality tests; it also deals effectively with diagnostic problems and therefore leads to more efficient and reliable estimates.

Table 9. The long run coefficients

| ARDL | | DOLS | |
|---------------------------------|----------------------|-----------------|------------------|
| Panel A: Long run elasticities | | | |
| Constant | .0429*** (.0414) | Constant | .0661*** (.0038) |
| G_t | .7934*** (.0485) | G_t | .8015*** (.0437) |
| $D1_t$ | .0222* (.0127) | $D1_t$ | .0184 (.0118) |
| $D2_t$ | .0888*** (.0165) | $D2_t$ | .0889*** (.0155) |
| Panel B: Hansen stability tests | | | |
| L_c | Test statistic .0643 | <i>p</i> Values | >.2 |

Note: The lags are chosen based on Schwarz information criterion (SIC). Figures in brackets are standard errors.

***Rejection of hypothesis at the 1% significance level.

Table 10. Error correction models and Granger causality tests

| Parameters | Dependent variable: ΔR_t | Parameters | Dependent variable: ΔG_t |
|---------------------------------|----------------------------------|-----------------------|----------------------------------|
| Panel A: Short run elasticities | | | |
| ϖ_{11} | – | ϖ_{21} | 1.090*** (.0560) |
| λ_{11} | .8053*** (.0414) | λ_{21} | – |
| θ_{11} | .0300* (.0163) | θ_{31} | –.0122(.0194) |
| θ_{21} | .1009***(.0153) | θ_{41} | –.1137***(.0178) |
| η_1 | –.6496***(.1356) | η_2 | –.6360***(.1266) |
| $F_{4,48}$ | 74.911*** | $F_{4,48}$ | 76.588*** |
| <i>R</i> ² | .90 | <i>R</i> ² | .91 |
| Panel B: Diagnostic tests | | | |
| $\chi^2_{Norm}(1)$ | 2.708 (.2582) | $\chi^2_{Norm}(1)$ | 1.042 (.5941) |
| $\chi^2_{LM}(1)$ | 2.060(.1512) | $\chi^2_{LM}(1)$ | .0655(.7979) |
| $\chi^2_{WT}(1)$ | 22.958(.1922) | $\chi^2_{WT}(1)$ | 23.431(.1746) |

Notes: χ^2_{Norm} is the Jarque–Bera test for normality, χ^2_{LM} is the Serial Correlation LM Test, and χ^2_{WT} is the white hetroskedasticity test. The lags used to estimate the ARDL model are chosen based on Akaike Information criterion (AIC). Figures in brackets are standard errors.

***Rejection of hypothesis at the 1% significance level.

Table 11. Tests for symmetric cointegration

| Parameters | Models | |
|---|----------------|----------------|
| | TAR | MTAR |
| ρ_1^a | -.8480 (.2239) | -.9734 (.2217) |
| ρ_2^a | -.7045 (.1933) | -.6112 (.1934) |
| Threshold (τ) | .0082 | .0107 |
| No cointegration $F(\rho_1 = \rho_2 = 0)$ b | 11.869*** | 12.958*** |
| Symmetry $F(\rho_1 = \rho_2)$ c | .02812 | 1.757 |
| k | 1 | 1 |
| Ljung-Box Q-statistics | .5290(.467) | 1.869(.172) |
| JB | 2.945(.2294) | 4.350(.1136) |

Note: "a" Entries in parentheses are the t-statistics for the null hypothesis of the ρ_i values. Entries of "b" represent the F-statistics that follow a nonstandard distribution of the sample values for the null hypothesis $\rho_1 = \rho_2 = 0$, and "c" Entries represent the sample F-statistics for the null hypothesis of symmetric adjustment $\rho_1 = \rho_2$. The appropriate lags for the TAR and M-TAR adjustment processes were chosen by AIC. Figures in brackets are standard errors. JB is the Jarque-Bera normality test which has an asymptotic $\chi^2(2)$ distribution.

***Significance level at 1%.

It is important to note however that evidence of fiscal synchronization does not necessarily imply that the budget deficit of the country in question is strongly sustainable. Rather, strong sustainability requires that government does not face difficulty marketing its debt (financing of public projects should not persistently exceed revenues). Evidently, the estimates of the long run elasticities presented in Table 9 suggest that the Nigerian fiscal deficit is weakly sustainable. The estimates of the adjustment parameters, indicating that a deviation from a balanced budget in one year is corrected by about 65 and 64% (Table 10), respectively, in the next year, which further suggests that the adjustment process toward equilibrium will take a long time, reinforce the existence of weak sustainability of budget deficits. Thus, it is likely that the finding of fiscal synchronization in both the short and long run implies that financing the budget deficit in the country (especially whenever there is revenue shortfall or falling crude oil prices) has been through increasing government debt.

The test results for the possible nonlinearity in the cointegration relationship between revenues and spending using Enders and Siklos (2001) cointegration techniques as discussed above are presented in Table 11.

The point estimates for ρ_1 and ρ_2 of both TAR and M-TAR models show convergence and the null hypothesis of no cointegration is rejected at the conventional significance levels. However, the null hypothesis of symmetric cointegration could not be rejected at all levels of significance. Thus, long run cointegration fails when asymmetric adjustment between the government revenues and spending is assumed. These results provide additional support for the robustness of the results obtained with the Johansen et al. (2000) and ARDL techniques.

5. Concluding remarks

The historical volatility of oil prices and that of revenue flows and expenditure in Nigeria may have a far-reaching effect on its budgetary process and the outlook of its fiscal prudence. Government revenues and expenditures can play an important role in managing federal budget deficits. This paper is an attempt to explore how government revenues and spending adjust to contain the size of the budget deficit in Nigeria as well as the sustainability of such deficit in the long run over the period 1961–2014 using error correction modeling framework.

The empirical evidence indicates that there is a long run relationship between revenues and spending after allowing for structural breaks. The results from both TAR and MTAR cointegration

models fail to support the hypothesis that the budgetary adjustment process in Nigeria is asymmetric. The policy implication of the finding that the size of the long run slope parameter is significantly less than one, which also offers support for the SBC strategy, suggests that the undiscounted value of government debts may not likely converge to zero in the limit. In other words, it shows the government's inability to redeem its debts from current bondholders at maturity without, for example, resorting to seigniorage, debt rescheduling, and/or surprise inflation, as such fiscal sustainability might not be easier to achieve in the future.

The Granger causality results indicate bidirectional short run causality between government revenues and spending through the partial *F*-statistics. For long run Granger causality, based on the error correction coefficients, the study also found evidence of a bidirectional long run causal relationship between government revenues and spending. These Granger causality results provide support for the fiscal, synchronization hypothesis in both short run and long run, indicating that the fiscal readjustment occurs simultaneously. It should be noted, however, that simultaneous adjustment here does not imply that the sustainability is strong, as mentioned above.

Finally, the fiscal sustainability evidence provided by this study (albeit weakly) seems to indicate that Nigeria still has chances of transforming its economic structure from oil dependent into a more diversified economy. Therefore, policy-makers should devise effective policies that will promote investment of the revenue windfall from crude oil sales (i.e. fiscal surpluses or foreign exchange reserve) into capital markets. Such investments will likely improve financial flows of the government, which in turn may guarantee financing capital projects and/or nonoil budget deficit(s). Policies that will ensure interests from the acquired financial assets are used to finance infrastructure development and are specifically desirable to make the economy more diversified.

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Notes

1. Johansen–Juselius technique requires that the linear combination of government revenues and expenditure (in the case of the present analysis) should produce *I*(0).
2. For details on the critical values for these tests, see Giles and Godwin (2012).
3. This methodology has been found to perform well, irrespective of whether the series are integrated of order zero or one (see, e.g. Beckmann, Belke, & Kühl, 2011; Kejriwal & Perron, 2008; Sharma & Setia, 2015 among others).
4. For details on the estimators of the variance matrices, see Bai and Perron (2003).

5. Note that the robustness checks will incorporate any possible structural break in the long run relationship between the variables.
6. For a discussion on this methodology, see Enders and Siklos (2001).
7. The result which indicates budget deficit is integrated of order one, *I*(1), was however expected.
8. Note that these percentages are calculated from the same data used for the present analysis.
9. To perform the ARDL cointegration test, two dummy variables were incorporated in Equation (13) as follows:

$$\Delta R_t = \delta_0 + \psi_R R_{t-1} + \psi_G G_{t-1} + \theta_1 D1_t + \theta_2 D2_t + \sum_{i=1}^p \varpi_i \Delta R_{t-i} + \sum_{i=0}^p \lambda_i \Delta G_{t-i} + \theta_{3+i} \Delta D1_{t-i} + \theta_{4+i} \Delta D2_{t-i} + \varepsilon_{1t}$$

where Δ denotes the first difference operator, *D1* denotes a dummy variable for the first break date that takes the value of 0 if *t* < 1990 and 1 otherwise, and *D2* is a dummy variable for the second break date that takes the value of 0 if *t* < 2000 and 1 otherwise

10. Allowing for structural breaks, the long-run coefficients of Equation (5) are estimated using the ARDL technique as. $\ln R_t = \delta_0 + \sum_{i=1}^p \psi_{1i} \ln R_{t-i} + \sum_{i=0}^p \psi_{2i} \ln G_{t-i} + \theta_1 D1_t + \theta_2 D2_t + \varepsilon_{1t}$, and DOLS methodology as follows $y_t = \alpha_0 + \lambda'_t \beta + \sum_{i=-l_1}^{i=l_2} \varphi_i \Delta x_{t+i} + \theta_1 D1_t + \theta_2 D2_t + \varepsilon_{1t}$ respectively.

where *D1* denotes a dummy variable for the first break date that takes the value of 0 if *t* < 1990 and 1 otherwise, *D2* is a dummy variable for the second break date that takes the value of 0 if *t* < 2000 and 1 otherwise.

11. The short run dynamics and Granger causality test within an error correction model are estimated as follows.

$$\begin{aligned}\Delta R_t = & \delta_{10} + \sum_{i=1}^p \omega_{1i} \Delta R_{t-i} + \sum_{i=1}^p \lambda_{1i} \Delta G_{t-i} + \theta_{1i} \Delta D_{t-i} \\ & + \theta_{2i} \Delta D_{t-i} + \eta_1 \hat{\varepsilon}_{t-1} + v_{1t}; \quad \Delta G_t = \delta_{20} \\ & + \sum_{i=1}^p \omega_{2i} \Delta G_{t-i} + \sum_{i=1}^p \lambda_{2i} \Delta R_{t-i} + \theta_{3i} \Delta D_{t-i} + \theta_{4i} \Delta D_{t-i} + \eta_2 \hat{\varepsilon}_{t-1} + v_{2t}\end{aligned}$$

where the dummy variables are as specified in Notes 9 and 10.

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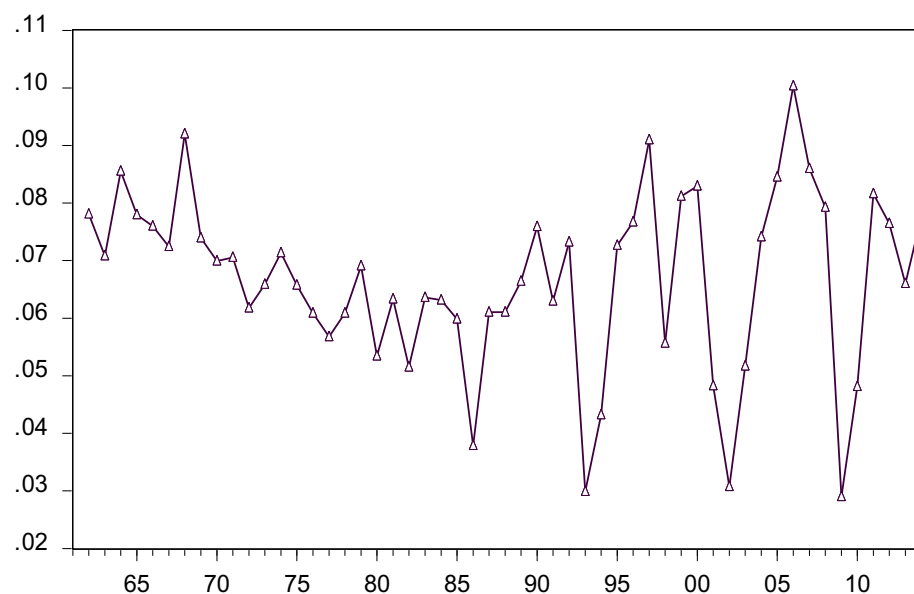
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Appendix A

Figure A1. Cointegration graph, 1961–2014.



Appendix B

Figure B1. Plots for revenues equation.

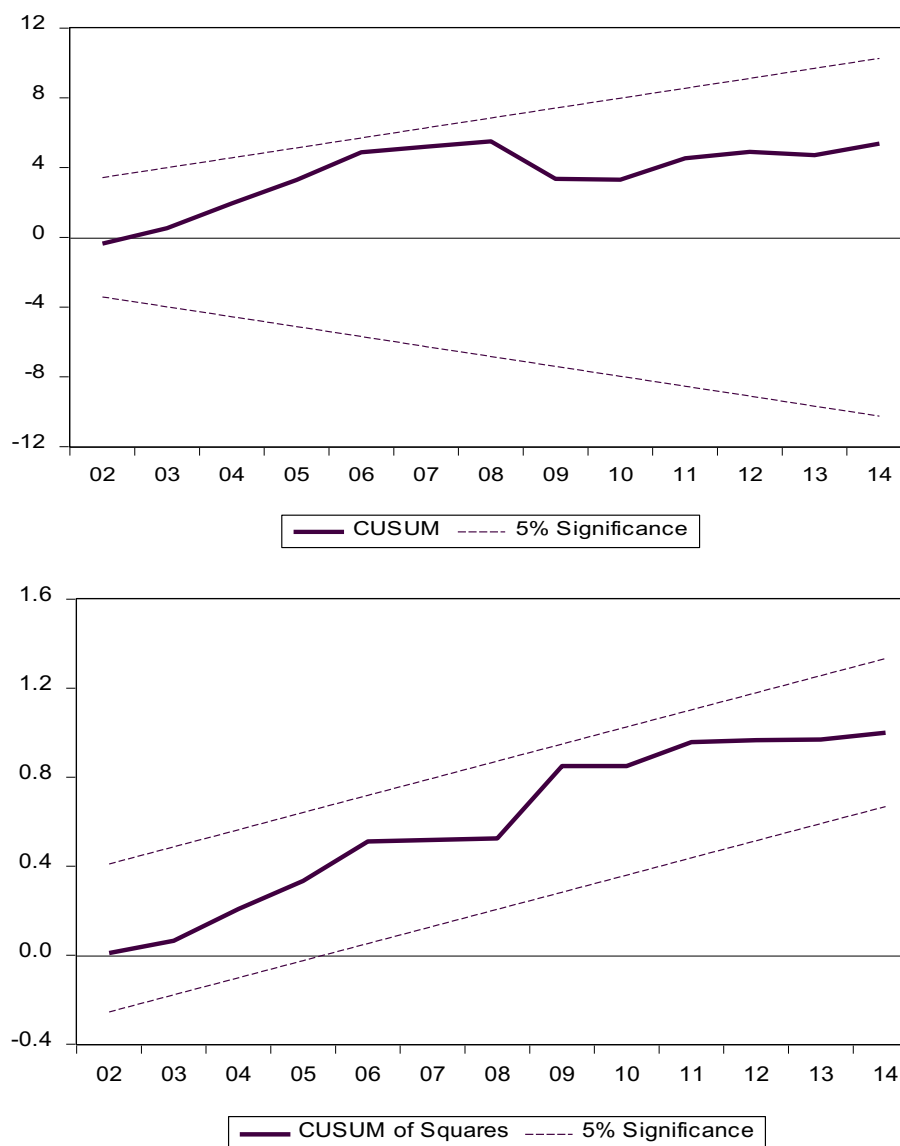
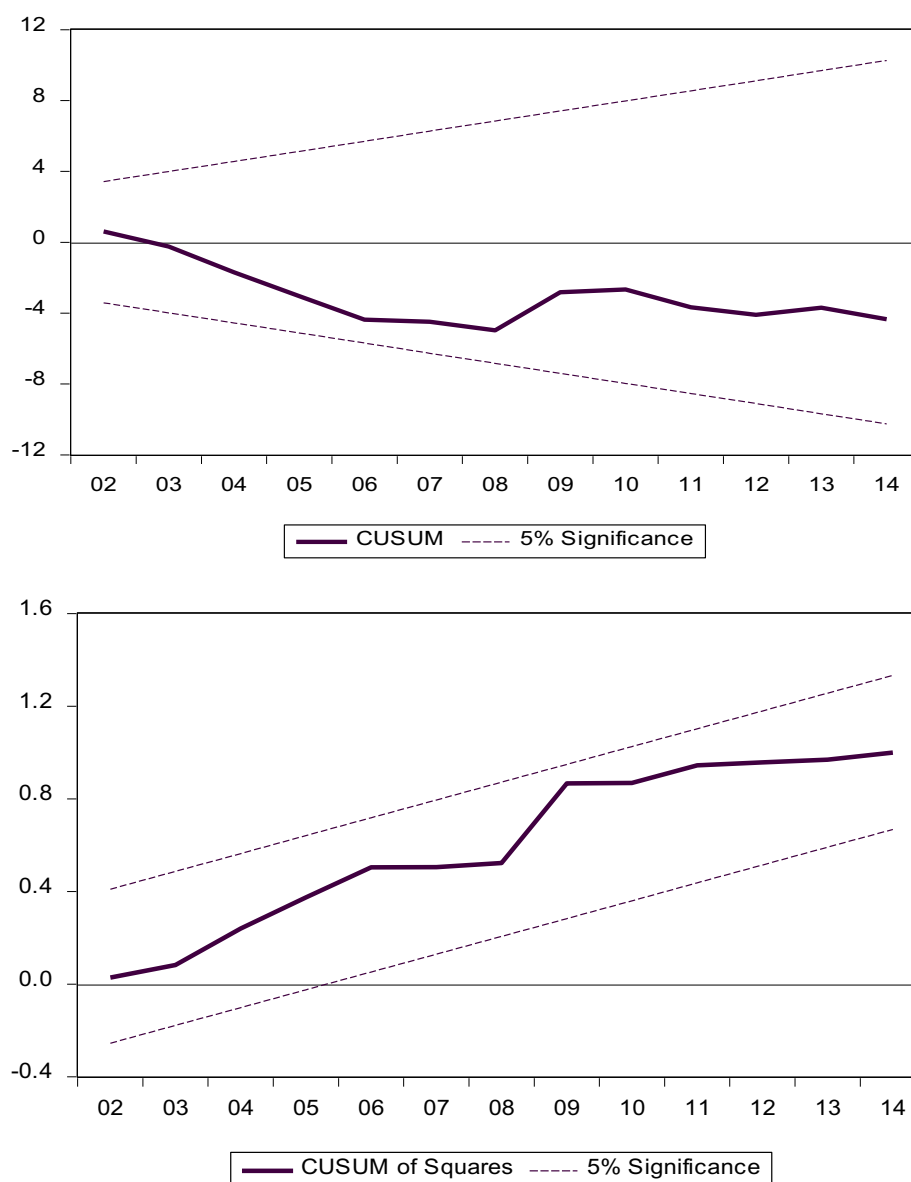


Figure B2. Plots for expenditures equation.



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