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The Energy Consumption-Growth Nexus in Seven Sub-Saharan African Countries

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Abstract

The paper investigates the long-run and the causality relationship between energy consumption and economic growth for seven Sub-Saharan African countries during the period 1970-2007. Using the bounds testing approach to cointegration, we find that energy consumption is cointegrated with economic growth in Cameroon, Congo, Cote d'Ivoire, and South Africa. Moreover, this test suggests that economic growth has a significant positive long run impact on energy consumption in these countries. Furthermore, causality tests suggest bidirectional causality between energy consumption and real GDP in Cote d'Ivoire and unidirectional causality running from real GDP to energy usage in the case of Congo.

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1. Introduction

Production and many consumption activities involve energy as an essential factor input in modern economies. It appears to be the key source of economic growth, industrialization and urbanization. Conversely, these latter variables may induce use of more energy, particularly commercial energy. Over the past few years, the relationship between energy consumption and economic growth has been extensively investigated. However, there seems to be no consensus about the relationship and the direction of causality between energy usage and economic development. Moreover, four hypotheses have been formulated to explain the direction of causality between energy consumption and real gross domestic product: growth, conservation, feedback, and neutrality hypotheses (Apergis and Payne 2009a, 2009b). The growth hypothesis implies that energy consumption contributes directly to economic growth within the production process as a direct input. Unidirectional Granger-causality from energy consumption to real GDP is consistent with the growth hypothesis; energy conservation policies could possibly reduce real GDP. The conservation hypothesis asserts that energy conservation policies designed to reduce energy consumption and waste may not have an adverse impact on real GDP. Granger-causality running from real GDP to energy consumption confirms the conservation hypothesis. The feedback hypothesis suggests that energy consumption and real GDP are interrelated and may serve as complements to each other. Hence, the existence of bidirectional Granger-causality between energy consumption and real GDP would substantiate the feedback hypothesis. Finally, the neutrality hypothesis considers energy consumption a relatively minor component of overall output and thus may have little or no impact on real GDP. In such cases, energy conservation policies may not adversely impact real GDP. The absence of Granger-causality between energy consumption and real GDP is evidence in favor of the neutrality hypothesis. As pointed out by recent studies, for instance Ferguson et al. (2000), Toman and Jemelkova (2003), Arbex and Perobelli (2010), the absence of any clear consensus on the relationship between energy consumption and growth can be attributed to the heterogeneity in climate conditions, varying energy consumption patterns, the structure and stages of economic development within a country, the alternative econometric methodologies employed, the presence of omitted variable bias along with varying time horizons of the studies conducted.

The general observation from literature is that most studies on the long-run and causal relationship between energy consumption and economic growth have been focused on developed economies. Not many studies have been reported in the case of Sub-Saharan Africa (Akinlo 2008, 2009, Wolde-Rufael 2009, and Keho 2007, among others). This study contributes to fill the gap. This paper aims at investigating the long-run and the causal relationship between energy consumption and economic growth in Cameroon, Congo, Cote d'Ivoire, Ghana, Kenya, Nigeria and South Africa over the period 1970-2007. Data series are from the African Development Bank (2008) and the World Bank (2008). We use two econometric approaches which have superior statistical properties in small samples: the Pesaran et al. (2001) bounds test and the Toda and Yamamoto (1995) version of Granger-causality testing procedure. In ought to test for cointegration between energy use and real GDP, we compute 'exact' bounds critical values based on stochastic simulations. Hence, we show that energy consumption is cointegrated with economic growth in Cameroon, Congo, Cote d'Ivoire, and South Africa. Moreover, this test suggests that economic growth has a significant positive long run impact on energy consumption in these countries. Furthermore, causality tests suggest bidirectional causality between energy consumption and real GDP in Cote d'Ivoire and unidirectional causality running from real GDP to energy usage in the case of Congo.

The remainder of this paper is organized as follows. Section 2 highlights the econometric framework. In the Section 3, we present the main results of this study. We finish by the conclusion.

2. Econometric framework

This section presents the econometric model used to study cointegration and causality between economic growth and energy consumption.

2.1 The cointegration approach

Econometric literature proposes different methodological alternatives to empirically analyse the long-run relationships and dynamic interactions between two or more time-series variables. The most widely used methods include the two-step procedure of Engle and Granger (1987) and the full information maximum likelihood-based approach due to Johansen and Juselius (1990). All these methods require that the variables under investigation are integrated of order one. This inevitably involves a step of stationary pre-testing, thus introducing a certain degree of uncertainty into the analysis. In addition, these tests suffer from low power and do not have good small sample properties (Cheung and Lai 1993, and Harris 1995). Due to these problems, this study makes use of the Pesaran et al. (2001) cointegrating test procedure.

The bounds testing approach to cointegration was originally introduced by Pesaran and Shin (1999) and further extended by Pesaran et al. (2001). The bounds testing approach to cointegration has at least two major advantages over the Johansen and Juselius (1990) approach. The first advantage is that it is applicable irrespective of whether the underlying regressors are purely $I(0)$, purely $I(1)$ or mutually cointegrated. The second advantage is that it has superior statistical properties in small samples. The bounds test is relatively more efficient in small sample data sizes as is the case in most empirical studies on African countries. Estimates derived from Johansen-Juselius method of cointegration are not robust when subjected to small sample sizes such as that in the present study.

To search for possible long run relationship amongst the variables, namely real gross domestic product (Y) and energy consumption (E), we employ the bounds testing approach to cointegration suggested by Pesaran et al. (2001). This involves estimating the following unrestricted error correction model (UECM):

$$\Delta \ln (Y_t) = \alpha_0+\alpha_1 t+\alpha_2 \ln (Y_{t-1})+\alpha_3 \ln (E_{t-1})+\alpha_4 DU_t+\sum_{i=1}^{m}\beta_i \Delta \ln (Y_{t-i})+\sum_{i=0}^{m}\gamma_i \Delta \ln (E_{t-i})+\varepsilon_t \tag{1}$$

where the $\alpha_i s$ $(i = 0, 1, 2, 3, 4)$, $\beta_i s$ $(i = 1, 2, ..., m)$ and $\gamma_i s$ $(i = 0, 1, 2, ..., m)$ are the parameters of the model. DU_t denotes a dummy variable capturing the period of structural adjustment in most of the Sub-Saharan African countries. $DU_t = 1$ after 1979 and zero otherwise. The structural lags m are determined by using minimum Akaike and Schwarz

Bayesian information criteria. To depict the presence of cointegration the estimated coefficients of lagged level variables are restricted equal to zero. Thus the null hypothesis for no cointegration between real GDP and energy consumption according to equation (1) is:

$$H_0 : \alpha_2 = \alpha_3 = 0 \tag{2}$$

The F-test statistic has a non-standard distribution which depends upon (*i*) whether variables included in the autoregressive distributed lags (ARDL) model are $I(0)$ or $I(1)$, (*ii*) the number of regressors, (*iii*) whether the ARDL model contains an intercept and/or a trend, and (iv) the sample size. Thus, the computed F-statistic is compared with two asymptotic bounds critical values tabulated by Pesaran et al. (2001). However, critical values reported by Pesaran et al. (2001) are generated for sample sizes of 500 observations and 1000 observations, with 20,000 and 40,000 replications, respectively. Given the relatively small sample sizes in our study (37 observations) we calculate critical values specific to our sample sizes. To this end, we generate the original set of critical values. These critical values are computed using stochastic simulations for a sample size T =37, based on 30,000 replications of the F-statistic used for testing the null of no cointegration in two models, one with an intercept but no trend and another one with both intercept and trend. Following Pesaran et al. (2001) notations, a model with an intercept and no trend is referred to as Case *III*, while a model with an intercept and an unrestricted trend is referred to as Case *V*, and is expressed as:

$$\Delta y_t = \xi_0 + \xi_1 t + \xi_2 y_{t-1} + \xi_3 x_{t-1} + \eta_t \tag{3}$$

Here, $t = 1, 2, \ldots, T$. Following Pesaran et al. (2001) notations, we have $z_{t-1} = (y_{t-1}, x_{t-1})'$, $w_t = (1, t)'$. The variables y_t and x_t are generated from $y_t = y_{t-1} + \eta_{1t}$ and $x_t = Px_{t-1} + \eta_{2t}$, with $y_0 = 0$, $x_0 = 0$ and $\eta_t = (\eta_{1t}, \eta_{2t})'$ is drawn as two independent standard normal variables. If x_t is purely $I(1)$, $P = 1$. On the other hand, $P = 0$ if x_t is purely $I(0)$. Two sets of critical values are generated. The lower critical value assumes that all the regressors are $I(0)$, while the upper critical value assumes that they are $I(1)$. Therefore, if the computed F-statistic is greater than the upper critical value, the null of no cointegration is rejected and we conclude that energy consumption and real GDP share a long-run level relationship. If the calculated F-statistic is below the lower critical value, then the null hypothesis of no cointegration cannot be rejected regardless of the orders of integration of the variables. On the other hand, if it falls inside the critical value band, the test is inconclusive unless we know the order of integration of the underlying variables.

If a cointegration relationship is observed between the series, Bardsen (1989) method will be used to estimate the short term ARDL model and compute the long-run coefficients. From the estimation of (1), the long-run coefficient is computed as the coefficient of the one lagged level explanatory variable divided by the coefficient of the one lagged level dependent variable and then multiplies with a negative sign. Thus, under the alternative of interest $\alpha_1 \neq 0$ and $\alpha_2 \neq 0$, the long-run level relationship between energy usage and growth is described by:

$$\ln(Y_t) = \vartheta_0 + \vartheta_1 t + \vartheta_2 DU_t + \vartheta_3 \ln(E_t) + \mu_t \tag{4}$$

where $\vartheta_0 = -\frac{\alpha_0}{\alpha_2}$, $\vartheta_1 = -\frac{\alpha_1}{\alpha_2}$, $\vartheta_2 = -\frac{\alpha_4}{\alpha_2}$, $\vartheta_3 = -\frac{\alpha_3}{\alpha_2}$ and μ_t is a stationary process with zero mean.

The existence of a cointegration derived from equation (1) does not necessarily imply that the estimated coefficients are stable. In this paper, we employ stability tests of Brown et al. (1975), which are also known as cumulative sum (CUSUM) and cumulative sum of squares (CUSUMSQ) tests based on the recursive regression residuals. These tests also incorporate the short-run dynamics to the long-run through residuals. The CUSUM and CUSUMSQ statistics are updated recursively and plotted against the break points of the model. Provided that the plots of these statistics fall inside the critical bounds of 5% significance, one assumes that the coefficients of a given regression are stable.

2.2 The causality approach

The Granger causality test is conventionally conducted by estimating vector autoregressive (VAR) models. Based upon the Granger representation theorem, Granger (1986) shows that if a pair of $I(1)$ series are cointegrated there must be a unidirectional causation in either way. If the series are not $I(1)$, or are integrated of different orders, no test for a long run relationship is usually carried out. However, given that unit root and cointegration tests have low power against the alternative, these tests can be inappropriate and can suffer from pre-testing bias. If the data are integrated but not cointegrated, then causality tests can be conducted by using the first differenced data to achieve stationarity. Granger non-causality test in an unrestricted VAR model can be simply conducted by testing whether some parameters are jointly zero, usually by a standard (Wald) F-test. Phillips and Toda (1993) show that the asymptotic distribution of the test in the unrestricted case involves nuisance parameters and nonstandard distributions. An alternative procedure to the estimation of an unrestricted VAR consists of transforming an estimated error correction model (ECM) into levels VAR form and then applying the Wald type test for linear restrictions. Toda and Yamamoto (1995) propose an interesting yet simple procedure requiring the estimation of an "augmented" VAR which guarantees the asymptotic distribution of the Wald statistic (an asymptotic χ^2-distribution), since the testing procedure is robust to the integration and cointegration properties of the process. We use a bivariate $VAR(p + d_{max})$ including real GDP (Y) and energy consumption (E), and examine the non-causality between these variables:

$$\ln(Y_t) = \psi_0 + \sum_{i=1}^{p} \psi_i \ln(Y_{t-i}) + \sum_{i=p+1}^{p+d_{\max}} \psi_i \ln(Y_{t-i}) + \sum_{i=1}^{p} \varphi_i \ln(E_{t-i}) + \sum_{i=p+1}^{p+d_{\max}} \varphi_i \ln(E_{t-i}) + \nu_{1t} \quad (5)$$

$$\ln(E_t) = \eta_0 + \sum_{i=1}^{p} \eta_i \ln(Y_{t-i}) + \sum_{i=p+1}^{p+d_{\max}} \eta_i \ln(Y_{t-i}) + \sum_{i=1}^{p} \lambda_i \ln(E_{t-i}) + \sum_{i=p+1}^{p+d_{\max}} \lambda_i \ln(E_{t-i}) + \nu_{2t} \quad (6)$$

$\psi_i s$, $\varphi_i s$, $\eta_i s$ and $\lambda_i s$ are the parameters of the model; $d_{\max}$ is the maximum order of integration suspected to occur in the system; $\nu_{1t} \sim i.i.d.N(0, \Sigma_{\nu 1})$ and $\nu_{2t} \sim i.i.d.N(0, \Sigma_{\nu 2})$ are the residuals of the model and $\Sigma_{\nu 1}$ and $\Sigma_{\nu 2}$ the covariance matrices of ν_{1t} and ν_{2t}, respectively. The null of non-causality from energy consumption to economic growth can be expressed as $H_0 : \varphi_i = 0, \forall\ i = 1, 2, ..., p$; $\varphi_i s$ $(i = 1, 2, ..., p)$ are the coefficients of the lagged values of $\ln(E)$ in the growth equation (5).

Let $\Phi = (\varphi_1, \varphi_2, \ldots, \varphi_p)$ be the vector of the first p VAR coefficients. For a suitable chosen R the Modified Wald Statistic for testing H_0 is computed using only the first p coefficients, as $W = T(vec(\hat{\Phi}))'R'[R\hat{\Sigma}R']^{-1}Rvec(\hat{\Phi})$; $\hat{\Phi}$ is the ordinary least squares estimate for the coefficient Φ and $\hat{\Sigma}$ is a consistent estimate for the asymptotic covariance matrix of $T^{1/2}vec(\hat{\Phi} - \Phi)$. The test statistic is asymptotically distributed as a χ^2 with p degrees of freedom. Two steps are involved with implementing the procedure: determination of (i) the lag length (p) and (ii) the maximum order of integration $(d_{\max})$ of the variables in the system of equations (5) and (6). In this study, we use the Akaike and Schwarz information criteria for the lag order selection. In addition, we employ the Augmented Dickey and Fuller (1979), the Phillips and Perron (1988) and the Ng and Perron (2001) unit root tests to determine the maximum order of integration.

3. Empirical results

This paper uses annual time series data on seven Sub-Saharan African countries, namely, Cameroon, Congo, Cote d'Ivoire, Ghana, Kenya, Nigeria and South Africa. The choice of countries included in this work was based on data availability. The data series comprise yearly observations between 1970 and 2007, and include real gross domestic product (GDP) as a measure for economic growth and energy usage in kilowatt per oil equivalent (Apergis and Payne, 2009a). Data are from two complement datasets: the 2008 World Development Indicators of the World Bank (2008) and the Selected Statistics on African Countries of the African Development Bank (2008).

While the unit root test is not a requirement for the bounds test for cointegration, it is important to establish that the variables are not integrated of an order higher than one. Another reason for conducing unit root tests is to determine the extra lags to be added to the vector autoregressive (VAR) model for the Toda and Yamamoto test. To ascertain the order of integration, we apply three tests for unit root, that are the Augmented Dickey-Fuller (1979)–ADF, the Phillips and Perron (1988)–PP and the Ng and Perron (2001)–NP– unit root tests. The results for the unit root tests about real GDP and energy consumption are summarized in Tables 1, 2 and 3. At the 5% level, the ADF, PP and NP tests provide strong evidence that the two series have a unit root for all the seven countries. They are integrated of order one.

Following the modelling approach described earlier, we determine the appropriate lag length and compute the bounds F-statistics. Models are estimated for m=0,1,..., 5. Table 4 provides results about the bounds tests F-statistics, and the lower and upper bounds critical values at 5% levels. It is shown that when real GDP is the dependent variable the computed F-statistics are less than the 5% lower bounds critical values. However, the computed F-statistics are higher than the 5% upper bounds critical values for four countries, namely, Cameroon, Congo, Cote d'Ivoire and South Africa, when the dependent variable is energy consumption. Hence, energy consumption and real GDP are cointegrated at the 0.05 significance level in these four countries. Given the findings reported in Table 4, we proceed with the empirical analysis only in the case of the countries where a long-run cointegrating relationship is established. Long-run effects of real GDP on energy consumption, and estimates for the dynamic relationship between these two variables are provided by Table 5. The computed lagged error correction term carries a significantly negative coefficient for the four countries. Moreover, the implementation of the Bardsen (1989) method suggests that real GDP has a significant positive long run impact on energy consumption. The long-run effect of real GDP on energy usage is relatively important in Congo (1.68), Cote d'Ivoire (1.15) and South Africa (1.05), but fewer in Cameroon (0.43). In the short run, economic growth impacts energy consumption positively in Cote d'Ivoire and South Africa but short-run fluctuations of real GDP seem to lower energy usage in Cameroon.

The CUSUM and CUSUMSQ stability tests were applied to the error-correction models and the graphs representing the tests are presented in Figures 1-7. It is shown that the plots of CUSUM and CUSUMSQ statistics are well within the critical bounds, implying that all coefficients in the error-correction models are stable. Therefore, this estimated models can be used for policy decision-making purposes.

As previously mentioned, to set the stage for the Toda-Yamamoto test, the order of integration of the variables is initially determined using the ADF, PP and NP unit root tests. Then, we determine the appropriate lag structures to include in the vector autoregressive models using Akaike and Schwarz Bayesian Information Criteria. Table 6 presents the results for the non-causality tests from energy consumption to economic growth, and vice versa, in the seven Sub-Saharan African countries. The third and sixth columns present the modified Wald statistics. It is shown that the null hypothesis that energy consumption does not Granger-cause economic growth in the short run has been rejected in favour of energy-led growth hypothesis in Cote d'Ivoire, at the 5% significance level. Similarly, there is enough evidence to support growth-led energy consumption in this country. This result suggests that there is a bidirectional relationship between energy consumption and economic growth in Cote d'Ivoire. This finding seems to support Odhiambo (2009) result for South Africa. The author found bidirectional causality between income and energy for South Africa using a trivariate error correction modeling. Our study also shows that there is a one-way causality running from economic growth to energy usage in Congo. The results show that a faster rate of growth promotes higher energy use. For most of the countries, economic growth does not depend significantly on energy use. Energy conservation policies may be implemented without adversely affecting economic growth. In fact, most of Sub-Saharan African countries depend largely on their agricultural sectors that are still the mainstay of their economies and provides employment for the majority of the population.

4. Conclusion and policy implication

This paper investigates the cointegration and causal relationship between energy consumption and economic growth in seven Sub-Saharan countries over the period 1970-2007. We make use of two recent econometric procedures appropriate for small sample data time series which are the Pesaran et al. (2001) bounds testing approach and the Toda and Yamamoto (1995) version of Granger-causality test. Data series are from two complementary sources: the 2008 selected statistics on African countries of the African Development Bank and the 2008 world development indicators of the World Bank.

It is found that energy consumption is cointegrated with economic growth in Cameroon, Congo, Cote d'Ivoire, and South Africa. Moreover, this test suggests that economic growth has a significant positive long run impact on energy consumption in these countries. Moreover, Granger-causality tests suggest bidirectional causality between energy consumption and real GDP in Cote d'Ivoire and unidirectional causality running from real GDP to energy usage in the case of Congo. These results suggest that, despite the large energy potential Africa has energy resources use remains very limited, especially in the agricultural sectors that dominate most of the economies in this area.

From a policy perspective, the results in this study are consistent with the feedback hypothesis that energy consumption and economic growth are interrelated and may very well serve as complements to each other, in Cote d'Ivoire. This suggests that energy consumption plays an important role in the growth prospects. Hence, Government of Cote d'Ivoire must ensure regular growth of energy production to boost the growth of output. In order to increase the production of energy, Cote d'Ivoire needs to rehabilitate and replace the existing infrastructure and facilities which are aging, inefficient and overloaded, and lead to problems of blackout and constant interruptions. Moreover, increasing the efficiency of current supply and utilization can help to develop power sector. Cote d'Ivoire should investigate and explore the possibilities of renewable energy for electricity generation to ensure uninterrupted energy supply. Furthermore, there is the need to ensure efficient energy use by the consumers. The consumers should be made aware of the importance of efficient use of electricity. Hence, Government should continue to educate the population about the efficient use of energy in households throughout the media. In Congo, economic growth Granger-causes energy use. Hence, Governement of Congo should promote economic growth.

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Countries	Growth			Energy		
	Lag	ln(Y)	Δln(Y)	Lag	ln(E)	Δln(E)
Cameroon	$4^{\dagger}$	-2.89 (-3.52)	-5.13*(-3.53)	0	-2.12 (-2.94)	-5.13*(-2.95)
Congo	1	-1.65 (-2.93)	-4.09*(-2.92)	1	-2.08 (-2.95)	-9.62*(-2.95)
Cote d'Ivoire	$1^{\dagger}$	-2.30 (-3.51)	-5.29*(-3.51)	$0^{\dagger}$	-2.85 (-3.54)	-6.53*(-3.54)
Ghana	1	-1.79 (-2.93)	-4.95*(-2.93)	5	-2.61 (-2.96)	-6.17*(-2.95)
Kenya	0	-2.31 (-2.95)	-6.03*(-2.95)	$0^{\dagger}$	-2.46 (-3.54)	-6.25*(-3.54)
Nigeria	0	-2.60 (-2.93)	-4.81*(-2.93)	0	-2.74 (-2.95)	-5.70*(-2.95)
South Africa	$1^{\dagger}$	-3.17 (-3.51)	-4.25*(-3.51)	$0^{\dagger}$	-2.72 (-3.54)	-4.87*(-3.54)

Notes: †Linear trend in model. 5% critival values in ( ). *denotes rejection of the null hypothesis of unit root at 5%. Δ indicates the first difference.

Table 1: ADF unit root tests results

Countries	Growth		Energy	
	ln(Y)	Δln(Y)	ln(E)	Δln(E)
Cameroon	-1.56 (-3.51)	-4.94*(-3.51)	-2.18 (-2.95)	-5.12*(-2.95)
Congo	-1.70 (-2.92)	-4.09*(-2.93)	-2.66 (-2.95)	-10.92*(-2.95)
Cote d'Ivoire	-2.39 (-3.51)	-5.37*(-3.51)	-2.83 (-3.54)	-6.52*(-3.54)
Ghana	-1.16 (-2.92)	-4.93*(-2.93)	-2.17 (-2.94)	-6.18*(-2.95)
Kenya	-1.91 (-2.92)	-7.01*(-2.93)	-2.39 (-3.54)	-8.15*(-3.54)
Nigeria	-1.95 (-2.92)	-4.87*(-2.93)	-2.86 (-2.96)	-8.17*(-2.95)
South Africa	-2.59 (-3.51)	-4.26*(-3.51)	-3.06 (-3.54)	-4.85*(-3.54)

Notes: 5% critival values in ( ). *denotes rejection of the null hypothesis of unit root at 5%. Δ indicates the first difference.

Table 2: Phillips-Perron unit root test results.

Countries		Growth			Energy
	Statistics	ln(Y)	Δln(Y)	ln(E)	Δln(E)
Cameroon	Z_α	-16.13 (-17.30)	-42.20*(-17.30)	-3.95 (-8.10)	-17.95*(-8.10)
	Z_t	-2.39 (-17.30)	-8.43*(-2.91)	-1.32 (-1.98)	-2.99*(-1.98)
Congo	Z_α	-2.41 (-8.10)	-16.43*(-8.10)	-2.97 (-8.10)	-14.09*(-8.10)
	Z_t	-1.02 (-1.98)	-2.86*(-1.98)	-1.15 (-1.98)	-2.65*(-1.98)
Cote d'Ivoire	Z_α	-2.75 (-17.30)	-22.28*(-17.30)	-6.02 (-17.30)	-17.65*(-17.30)
	Z_t	-1.13 (-2.91)	-3.32*(-2.91)	-1.71 (-2.91)	-2.97*(-2.91)
Ghana	Z_α	-4.06 (-8.10)	-21.70*(-8.10)	-3.70 (-8.10)	-13.73*(-8.10)
	Z_t	-1.37 (-1.98)	-3.29*(-1.98)	-1.03 (-1.98)	-5.18*(-1.98)
Kenya	Z_α	-0.17 (-8.10)	-9.54*(-8.10)	-5.65 (-17.30)	-17.48*(-17.30)
	Z_t	-2.59 (-1.98)	-2.10*(-1.98)	-1.56 (-2.91)	-2.95*(-2.91)
Nigeria	Z_α	-4.69 (-8.10)	-18.12*(-8.10)	-0.07 (-8.10)	-15.49*(-8.10)
	Z_t	-1.70 (-1.98)	-2.91*(-1.98)	-0.05 (-1.98)	-2.77*(-1.98)
South Africa	Z_α	-3.68 (-17.30)	-18.62*(-17.30)	-2.44 (-17.30)	-17.49*(-17.30)
	Z_t	-1.35 (-2.91)	-3.05*(-2.91)	-0.97 (-2.91)	-2.92*(-2.91)

Notes: 5% critival values in ( ). *denotes rejection of the null hypothesis of unit root at 5%. Δ indicates the first difference.

Table 3: Ng-Perron unit root test results.

Countries	Dep. var.	Lag	$\chi^2(1)$	F-stat.	5% lower c. v.	5% upper c. v.	Cointegration
Cameroon	ln(Y)	1	0.060	2.51	5.407	6.312	No
	$\ln(E)^a$	2	0.031	12.05	7.324	8.192	Yes
Congo	ln(Y)	1	0.172	2.75	5.407	6.312	No
	$\ln(E)^a$	0	2.203	16.05	7.324	8.192	Yes
Cote d'Ivoire	ln(Y)	4	0.002	5.23	5.407	6.312	No
	$\ln(E)^a$	0	0.863	17.40	7.324	8.192	Yes
Ghana	$\ln(Y)^a$	4	2.219	1.89	7.324	8.192	No
	$\ln(E)^a$	4	0.002	7.23	7.324	8.192	No
Kenya	ln(Y)	1	2.467	3.33	5.407	6.312	No
	ln(E)	1	0.369	2.19	5.407	6.312	No
Nigeria	$\ln(Y)^a$	3	0.44	6.54	7.324	8.192	No
	$\ln(E)^a$	3	0.068	6.04	7.324	8.192	No
South Africa	$\ln(Y)^a$	0	0.940	1.43	7.324	8.192	No
	ln(E)	3	1.796	8.82	5.407	6.312	Yes

Notes: ^{a} Linear trend in equation. Dep. var. and c.v. are related to dependent variable and critical value. 5% critical values are calculated using stochastic simulation with 30,000 replications. $\chi^2(1)$is an LM statistic for testing no residual serial correlation against order 1.

Table 4: Bounds tests, F-statistics.

Independent variables	Cameroon $\Delta\ln(E)$	Congo $\Delta\ln(E)$	Cote d'Ivoire $\Delta\ln(E)$	South Africa $\Delta\ln(E)$
$Constant$	2.41*(4.28)	-6.98*(-4.22)	-2.77*(-3.34)	
$Trend$	0.001(1.66)	0.01**(2.43)	0.02*(4.66)	
DU_t	0.16*(2.71)		0.24*(4.38)	
$\ln(Y_{t-1})$	0.45*(4.28)	1.63*(5.63)	0.93*(5.38)	0.16*(5.93)
$\ln(E_{t-1})$	-1.04*(-5.47)	-0.97*(-6.97)	-0.81*(-5.71)	-0.15*(-5.88)
Long run effect	0.43*(6.40)	1.68*(4.71)	1.15*(10.35)	1.05*(21.8)
$EC(-1)^a$	-1.04*(-5.47)	-0.97*(-6.97)	-0.81*(-5.71)	-0.15*(-5.88)
$\Delta\ln(Y_t)$			1.17*(3.96)	0.46*(3.03)
$\Delta\ln(Y_{t-1})$	-0.43**(-2.15)			
$\Delta\ln(E_{t-1})$	0.63*(3.75)			
$\Delta\ln(Y_{t-2})$	-0.54**(-2.59)			
$\Delta\ln(E_{t-2})$	0.33 (1.90)			
$\Delta\ln(E_{t-3})$		-2.06*(3.28)		-0.36**(-2.32)
$R-squared$	0.60	0.63	0.67	0.61
$\chi^2(1)$ $[p-value]$	0.037[0.85]	2.202[0.14]	0.467[0.49]	0.366[0.54]

Notes: [a]EC(-1) denotes the coefficient estimate of the lagged error correction term.
*and **indicate significance at the 1% and 5% respectively. Numbers in ( ) are t-statistics.

Table 5: ARDL estimation.

Countries	ln(E) does not cause ln(Y)			ln(Y) does not cause ln(E)		
	Lag	M-Wald	P-value	Lag	M-Wald	P-value
Cameroon	2	1.64	0.441	2	0.21	0.901
Congo	2	0.79	0.674	2	8.76*	0.012
Cote d'Ivoire	1	4.43	0.044*	1	4.57*	0.032
Ghana	2	2.09	0.352	2	5.45	0.065
Kenya	2	2.18	0.336	2	0.12	0.943
Nigeria	4	1.54	0.672	4	4.70	0.195
South Africa	1	0.01	0.913	1	0.36	0.548

Notes: *indicates significance at the 5% significance level.

Table 6: Non-causality test results.

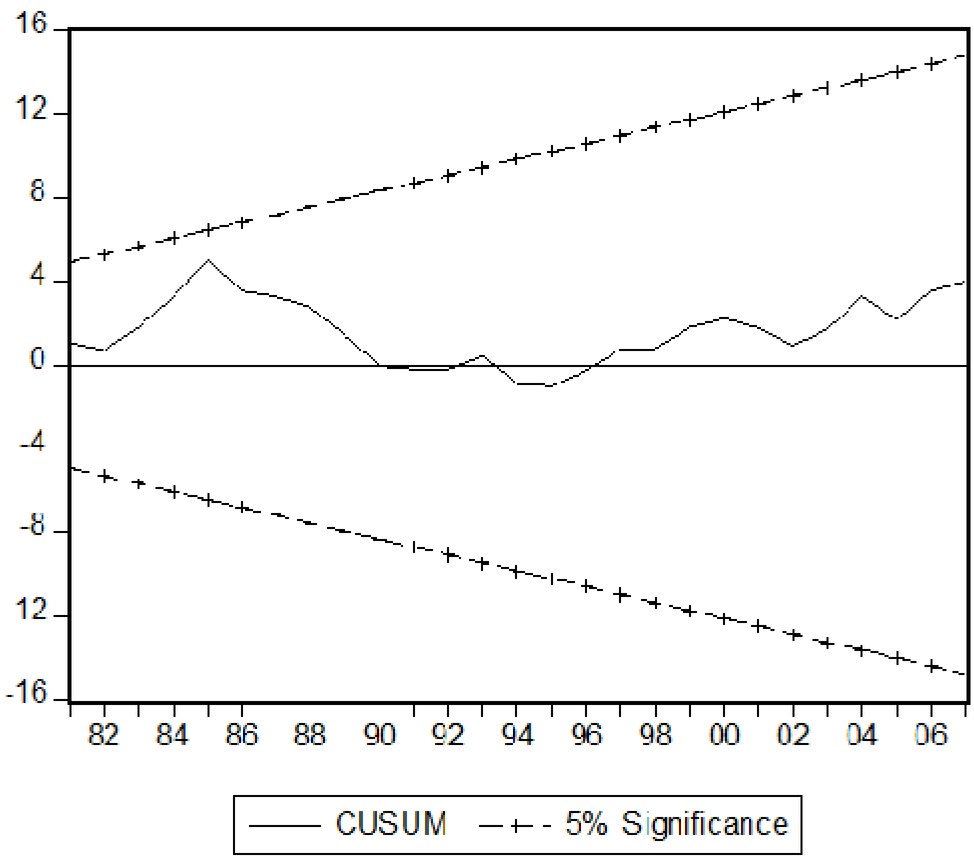

Figure 1: Plot of CUSUM, Cameroon.

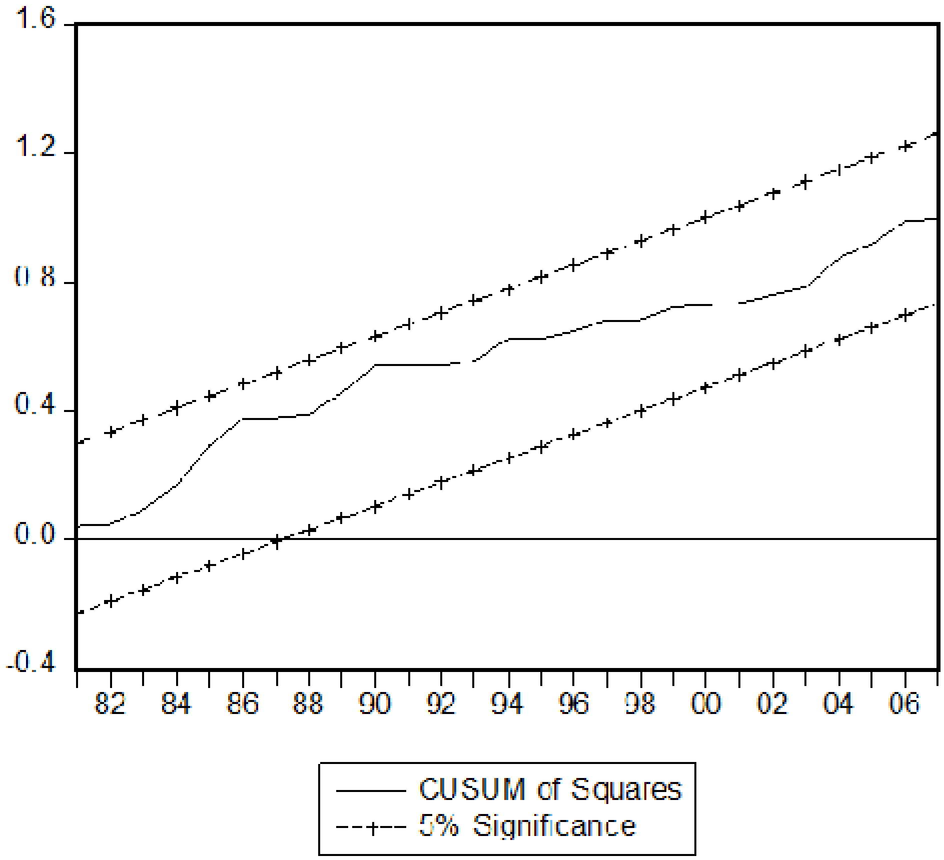

Figure 2: Plot of CUSUM of squares, Cameroon.

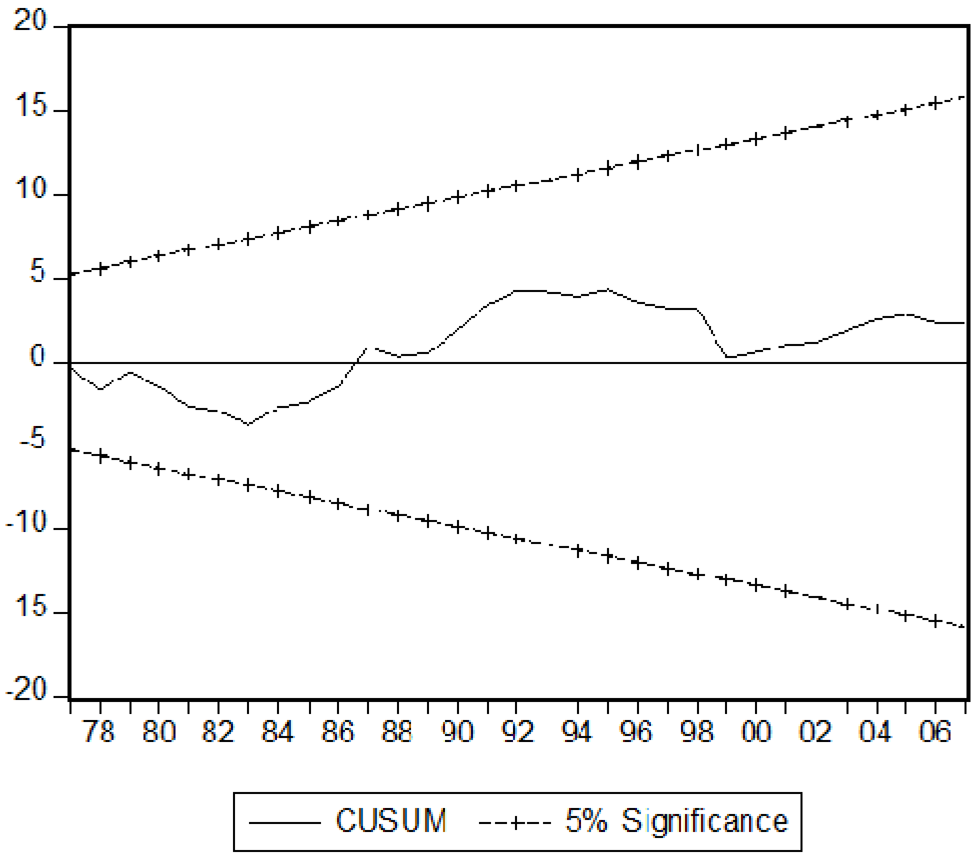

Figure 3: Plot of CUSUM, Congo.

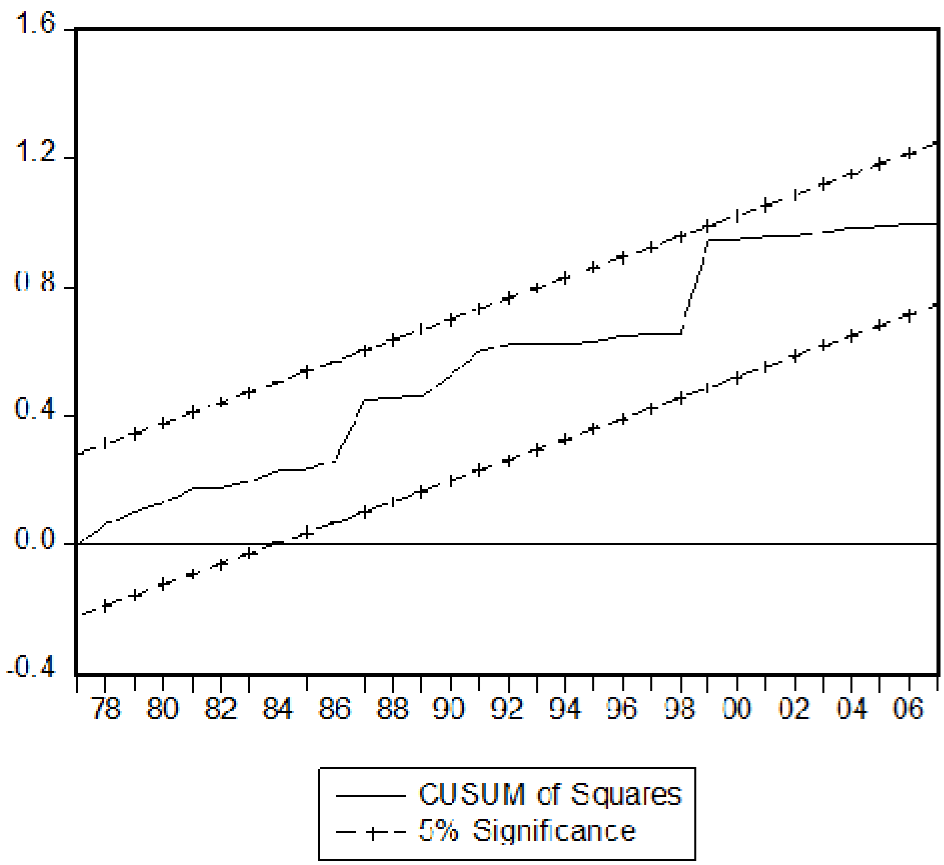

Figure 4: Plot of CUSUM of squares, Congo.

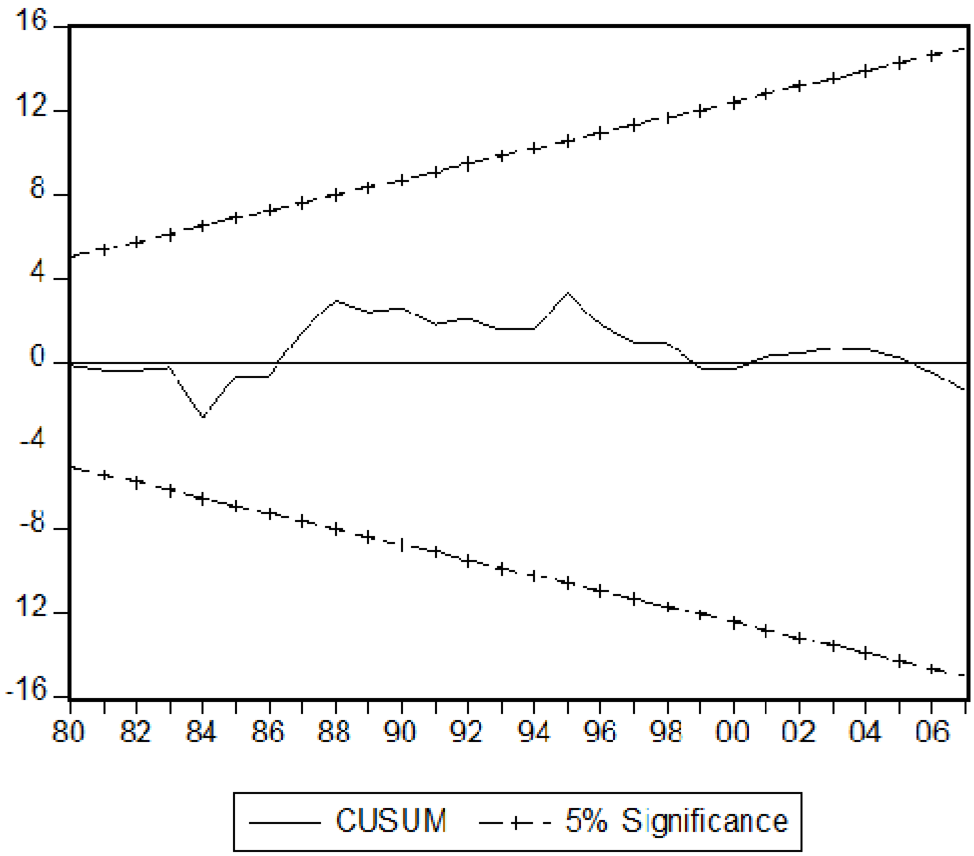

Figure 5: Plot of CUSUM, Cote d'Ivoire

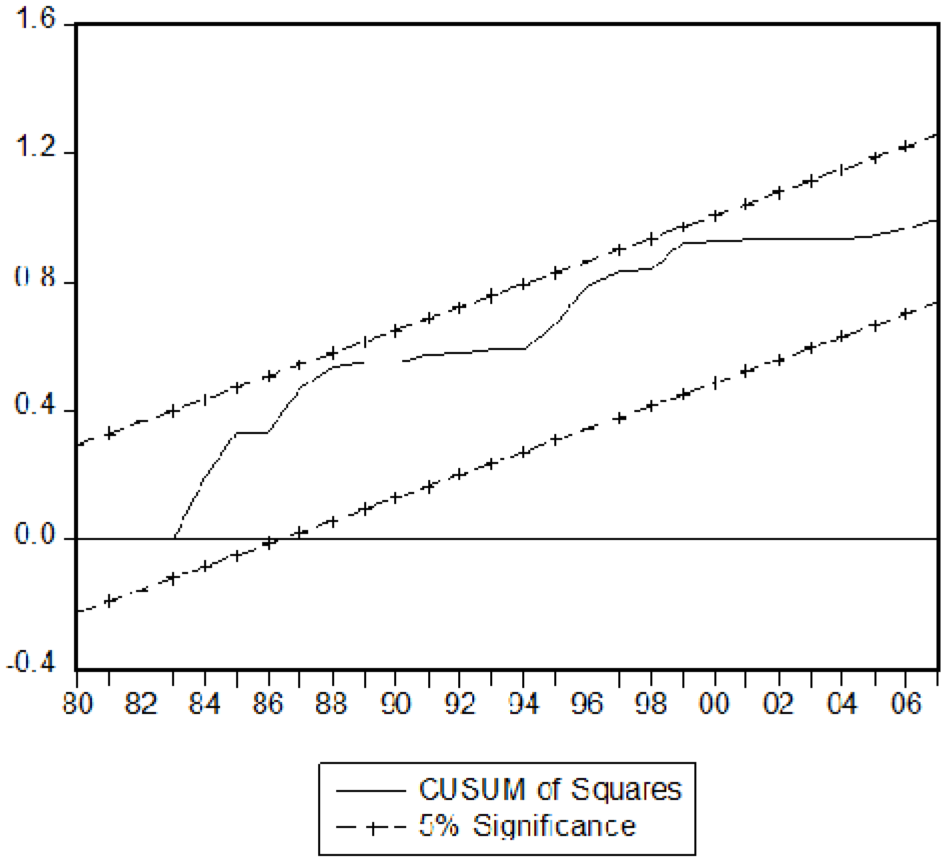

Figure 6: Plot of CUSUM squares, Cote d'Ivoire.

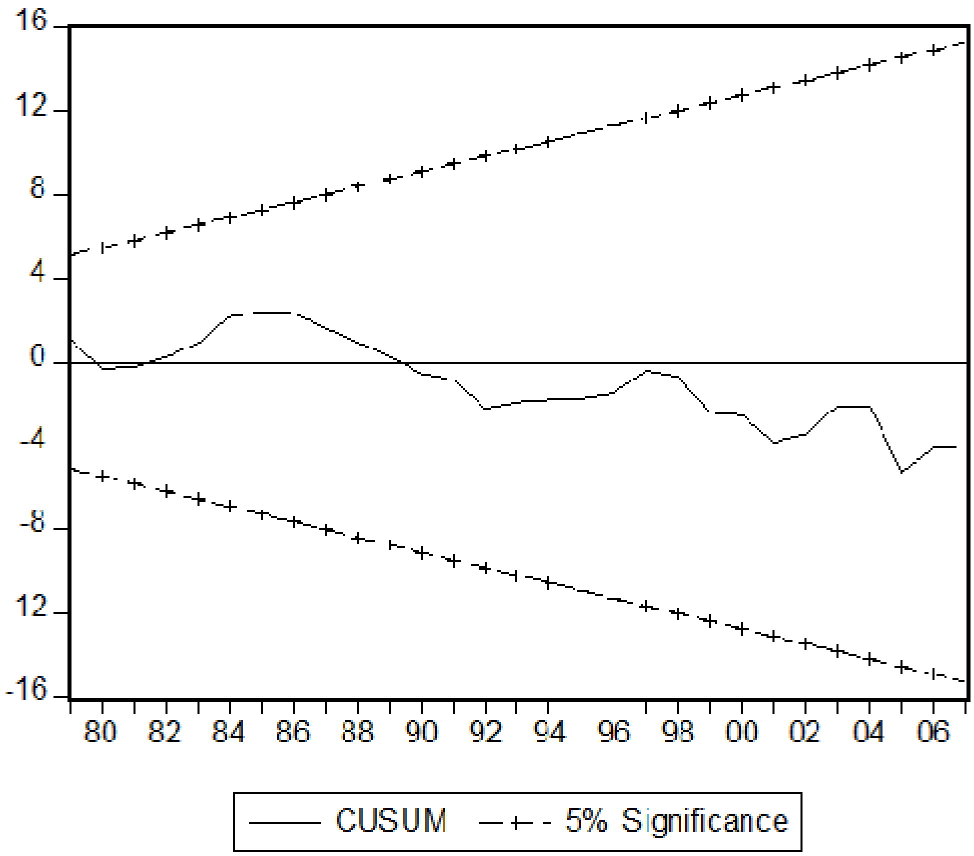

Figure 7: Plot of CUSUM, South Africa.