DYNAMIC LINKAGES AND GRANGER CAUSALITY BETWEEN TRADE AND BUDGET DEFICITS: EVIDENCE FROM AFRICA

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ABSTRACT
This paper provides new evidence on the long-run relationship between trade and budget deficits in ten African countries over the quarterly period 1973:2 - 2005:4. Cointegration analyses are based on four approaches: Harris-Inder (1994), Shin (1994), Geweke and Porter-Hudak (1983) and Sowell (1992). In conformity with theoretical considerations, our analysis reveals that there is a positive long-run relationship between the trade deficit and the budget deficit; however, in the short run, we find weak evidence that these deficits are closely linked and that the budget deficit causes the trade deficit. The analysis finds that bidirectional long-run causality between deficits receives strong empirical support. Unidirectional causality and no causality characterize the short run, so bidirectional causality was found to be largely unimportant. Budget deficit adjustment, not trade deficit adjustment, is shown to be the key engine governing the speed of budget-trade deficit convergence; that is, the budget deficit is the primary variable that changes in order to restore equilibrium when the system has been subjected to shock. Moreover, budget deficits are found to converge much faster than trade deficits.

Keywords: Budget deficit, trade deficit, cointregation and Africa.

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I. INTRODUCTION

One of the less heavily investigated subjects in the literature of developing countries is the relationship between trade and budget deficits. The main purpose of this paper is to empirically investigate this relationship for ten African countries over the quarterly period 1973:2 through 2005:4 utilizing the cointegration test procedures and the related notion of error-correction. The African nations examined are Botswana, Burundi, Kenya, Mauritius, Nigeria, Rwanda, Sierra Leone, South Africa, Togo and Tunisia.

An important contribution of this study is that it examines the relationship between the trade deficit and the budget deficit in the context of conventional and fractional cointegration approaches. Additionally, this paper uses trivariate causality to investigate empirically the causal relation between two deficit variables. These approaches have two major advantages over the extant empirical research on this issue. First, this paper relaxes the assumption in the conventional cointegration analyses that the cointegrating residual must be integrated of order zero (i.e., I(0)) before the null hypothesis of no-cointegration can be rejected. The strict nature of I(0) assumption may lead to invalid conclusions since it may not be adequate in assessing each country's data for the existence of a long-run equilibrium relationship.

In this study, we employ tests for fractional cointegration, and two estimators used are the GPH test of Geweke and Porter (1983) and the exact maximum likelihood (EML) of Sowell (1992) and others. This generalized form of cointegration enables us to capture slow or subtle mean-reverting dynamics because the integration order of the cointegrating residual is allowed to take any value on the real line. Therefore, it enables the capture of a long-run equilibrium relationship that is mean reverting though stationary but not exactly I(0).

Second, tests for cointegration and Granger causality are conducted here in a trivariate framework. It is well-known that studies which focus only on two variable cases may be biased due to the omission of an important third variable see Lutkepohl (1982, 1993). As Lutkepohl demonstrates, non-causality in a bivariate setting may be associated with the methodological problem of a third missing variable. Several researchers have shown that this omission-of-variables bias often found in a bivariate setting could result in the statistical instability of the cointegration space and, hence, distorts cointegration inferences, see Arize et al. (2000).

The objective of this paper is threefold. The first objective is to explore the plausible long-run linkage between the trade deficit and the budget deficit in 10 African countries. A large number of studies have applied the standard cointegration tests in the investigation of the issue, but very few apply the fractional and /or Harris-Inder (1994) as well as Shin (1994) tests. The three cointegration test methods applied here are based on two different null hypotheses: absence of cointegration in the fractional methods and presence of cointegration in the Harris-Inder or Shin method. To our knowledge, no other study provides an investigation of trade-budget-deficits relationship using African countries' data and fractional and /or Harris-Inder (1994) as well as Shin (1994) cointegration tests.

The second objective of this article is to examine Granger causality between trade and budget deficits within the well known cointegration and error-correction framework. This framework tries to establish causality between the two variables after reintroducing the low-frequency (long-run) information (through the error-correction term) into the analysis. Hence, we are able to test for long and short-run causality, as well as identify which variable(s) bear the burden of short-run adjustment to reestablish the long-run equilibrium following a shock to the system. From a policy-making perspective, these issues are important in order to eliminate trade and budget deficits; it is generally accepted that an economy without these deficits is more likely to thrive than an economy overwhelmed with deficits.

The third objective of this study is to estimate the speed of adjustments, median and mean time lags of trade and budget deficits in African countries. How quickly the trade deficit (the budget deficit) responds to changes in the budget deficit (the trade deficit) has not been examined in this literature. Understanding the adjustment process is likely to lead to improved predictions that can benefit both market participants and policymakers. For instance, knowledge of the primary variable that adjusts its level in order to restore equilibrium when the system has been subjected to shock can guide authorities in Africa in deciding the appropriate exchange rate, export promotion, productivity improvement, fiscal, and monetary policies to be implemented in adjusting macroeconomic variables such as the trade balance. Understanding the nature of the adjustment process for budget/trade deficits can also be useful in evaluating the future effects that may occur as a result of policy changes and can provide useful guidance in drawing conclusions regarding any general hypotheses. In sum, this article differs from the existing literature, not only in its data set, but also in its empirical method and its new evidence identifying budget deficit adjustment to be the key engine governing the speed of budget-trade deficit convergence.

The remainder of this paper is organized as follows. In Section II, we examine some of the simple theoretical foundations upon which our empirical results are based. The empirical results are presented in Section III while summary and concluding remarks are in Section IV.

II. THEORETICAL CONSIDERATIONS

The well-known saving-investment identity could be derived from the national income identity:

$$Y / C + I + G + (X - M); \text{ and } S + (T - G) / I + (X - M) \tag{1}$$

where Y is the national income; C is private consumption; I is investment; G is government expenditures on final goods and services; X is exports of goods and services; M is imports of goods and services; (X - M) is net exports or the trade balance; S is national savings (private sector savings; and T is government tax revenues. After substituting, equation (1) becomes

$$(X - M) / (S - I) + (T - G); \quad \text{or } TD = BD + SI. \qquad (2)$$

In this case, net exports simply equal the private saving-investment gap plus the budget balance. Thus, assuming a stable saving - investment gap, an increase in public sector deficit will directly increase the trade deficit. While the identity does not provide any behavioral or temporary relationships between the deficits, or even the direction of causality, it provides a basis for expecting a positive long-run equilibrium relationship between the two deficits.

Four possible causation linkages may be present between the budget deficit and the trade deficit. The first linkage is the Keynesian (conventional) proposition often associated with the Mundell-Fleming model. It argues that there exists a positive relationship between the two deficits and that causality is from the budget deficit to the trade deficit. In the context of this model, an increase in budget deficit would cause an increase in domestic interest rate above the world rate, with capital inflows and appreciation of the domestic currency as effects. These effects, in turn, result in an increase in trade (current account) deficit.

As discussed in Kearney and Monadjami (1990), reverse causation from trade to budget deficits can come about if there is a change in the expectations of inflation. A decrease in expected inflation would lead to currency appreciation and thus decrease net exports and increase the trade deficit. This in turn will have the usual multiplier-type decrease in output and consequently in tax revenues. Thus, by this approach, decreased inflationary expectations would lead trade deficits to cause budget deficits. Also, reverse causation from trade to budget deficits can occur if excessive trade deficits plunge an economy into a recession and subsequently lead to a financial or solvency crisis in which a large injection of public funds may be needed to rehabilitate the struggling financial sector or to minimize the severity of a recession. For more discussion of this, see Kim and Kim (2006). In other words, the large inflow of capital or debt accumulations affects the budgetary stance of a country and ultimately leads to budget deficit. Summers (1988) has referred to this reverse causality as current account targeting and suggests that external adjustments may be sought through budget (fiscal) policy.[1]

Bidirectional causality may exist between trade and budget deficits. While budget deficits may cause trade deficits, the existence of significant feedback may cause

[1] Examples of studies that support reverse causation are Kearney and Monadjami (1990), Anoruo and Ramchander (1998), Khalid and Teo (1999) and Alkswani (2000).

causality between the two variables to run in both directions. In this case, it is not enough to cut the budget deficit in order to eliminate trade deficits. It is necessary as well to complement budget-cut policies with a coherent package focusing on policies for export promotion, productivity improvement and exchange rate, among others.

In contrast, proponents of the Ricardian equivalence hypothesis (REH) suggest the absence of any relationship between the trade deficit and the budget deficit. Proponents of this view point out that, while a tax cut (hence a deficit) has the effect of reducing public revenues and public savings and enlarging the budget deficit, it increases private saving by an amount equal to the expected increase in the tax burden in future years. That is, savings will respond positively to the changes in budget deficits, leaving the trade deficit unaltered. Similarly, if government runs a deficit by borrowing, the economic agents expect that government will raise future taxes to finance the budget deficit and so they increase their savings to meet the future tax burden. In sum, alterations in the composition of public financing (i.e., debt versus taxes) have no impact on real interest rate, aggregate demand, private spending, the exchange rate or current account balance. In other words, the absence of any Granger causality relationship between the two deficits would support REH.

III. METHODOLOGY AND ESTIMATION RESULTS

A. The data and the unit root tests

The quarterly data (1973:2 - 2005:4) for our investigation come from the International Monetary Fund (IMF)'s International Financial Statistics (IFS) CD-ROM (May, 2005). When data are unavailable, other sources such as IFS CD-ROM (June 1998) and World Tables have been used to fill-in any data gaps. Trade deficits (excess of imports over exports) in U.S. Dollars were converted to domestic currency by using the relevant nominal exchange rates. To eliminate the upward bias that may result from increases in economic growth and inflation, we have scaled the trade deficits in domestic currency by the nominal gross domestic product (GDP). Henceforth, the trade deficit is referred to as TD_t. Similarly, budget deficits in domestic currency have been scaled by nominal GDP and will henceforth be referred to as BD_t. To account for different channels of interaction between the trade/deficit process, we have included industrial production (to control for the effects of business cycles) or real output as one of the right-hand side variables. In what follows, we refer to this as activity measure (y_t).

B. Cointegration test results

To establish whether there is a long-run equilibrium relationship among the variables TD, BD and y, we must employ the concept of cointegration developed by Engle and Granger (1987). The basic idea of cointegration is that two or more nonstationary time series may be regarded as defining a long-run equilibrium relationship if a linear combination of the variables in the model is stationary (converges to any equilibrium over time). Thus, if the link between the trade deficit, the budget deficit and the activity measure describes a stationary long-run relationship, this can be interpreted to mean that the stochastic trend in the variables is related. In other words, even though deviations from the equilibrium should occur, they are mean reverting.

A prerequisite in applying the cointegration procedure is to test the unit root properties of the series. The common practice is to use the augmented Dickey–Fuller (ADF) t-statistics. As is usually done in the literature, we report the value of the ADF (k), where k is the minimum lag for white errors. According to Table 1, the ADF test results suggest that the null hypothesis of a unit root (i.e., nonstationarity) is accepted when the variables are in levels, but it is rejected when the series are in first differences. Series that are nonstationary in levels have a unit root (stochastic trend). Shocks to a time series that has a unit root are, in part, permanent; they change the long-run level of the series permanently.

Table 1
Results of the ADF test for a Unit root tests

Name of Country	***Levels***			***Differences***		
	TD	***BD***	***y***	***TD***	***BD***	***y***
Botswana	***-2.01***	***-1.88***	***2.35***	***-5.56***	***-6.21***	***-4.14***
Burundi	***-1.80***	***-2.43***	***-1.72***	***-5.92***	***-7.24***	***-4.10***
Kenya	***-4.13***	***-5.01***	***-2.14***	***-6.25***	***-6.06***	***-8.25***
Mauritius	***-2.72***	***-2.00***	***-1.37***	***-5.28***	***-5.69***	***-5.00***
Nigeria	***-2.98***	***-2.50***	***-1.30***	***-5.31***	***-5.64***	***-5.36***
Rwanda	***-3.02***	***-2.65***	***-1.97***	***-6.41***	***-6.06***	***-5.05***
Sierra Leone	***-2.14***	***-2.65***	***-1.26***	***-5.35***	***-5.97***	***-4.89***
South Africa	***-3.19***	***-2.38***	***0.27***	***-5.66***	***-6.41***	***-3.47***
Togo	***-2.72***	***-6.63***	***-3.01***	***-5.38***	***-6.74***	***-4.79***
Tunisia	***-3.60***	***-3.33***	***1.09***	***-6.38***	***-5.66***	***-2.52***

Notes: - The augmented Dickey - Fuller (ADF) test examines the null hypothesis of an I(1) process against the alternative of an I(0) process. The number of lagged differences included in the ADF regression is eight and procedures in Arize and Ndubizu (1992) and Arize and Shwiff (1998) were followed. The critical value at the 5 percent level for the ADF tests is -3.45 and for the first difference it is -2.86.

C. Conventional cointegration tests

Since a unit root has been confirmed for the data series, the question is whether there exists some long-run equilibrium relationship among trade deficits, budget deficits and activity measure for each country in our sample. As can be seen in Table 2, the Harris-Inder test as well as the Shin's test is conducted with three, six and eight lags. The null of stationarity (or cointegration) is not rejected at all lag lengths in some countries. However, as the authors suggest, some degree of augmentation in the tests is needed for better results.

Table 2
Cointegration tests

Name of Country	Harris - Inder						Shin test						
	Lag series			Lag series			Lag series			Lag series			Estimator
	3	6	8	3	6	8	3	6	8	3	6	8	DLS
	TD as dependent			BD as dependent			TD as dependent			BD as dependent			ψt[t]
Botswana	0.21	0.16	0.14	0.20	0.18	0.17	0.06	0.07	0.08	0.05	0.05	0.06	1.43[4.77]
Burundi	1.15	0.81	0.69	0.18	0.19	0.19	0.85	0.62	0.53	0.13	0.13	0.13	0.47[1.94]
Kenya	0.07	0.06	0.05	0.16	0.14	0.14	0.08	0.06	0.06	0.08	0.07	0.08	0.41[1.85]
Mauritius	0.14	0.11	0.10	0.59	0.35	0.29	0.09	0.08	0.07	0.32	0.23	0.20	0.22[1.67]
Nigeria	0.13	0.10	0.09	0.46	0.32	0.27	0.06	0.05	0.05	0.35	0.25	0.22	0.13[1.60]
Rwanda	1.05	0.89	0.79	0.23	0.21	0.19	1.35	0.96	0.82	0.17	0.16	0.15	0.68[2.17]
Sierra-Leone	0.89	0.60	0.51	0.62	0.48	0.43	0.20	0.16	0.15	0.82	0.68	0.62	0.57[2.19]
South-Africa	0.12	0.08	0.07	0.20	0.14	0.12	0.09	0.07	0.06	0.21	0.15	0.13	0.58[3.18]
Togo	0.20	0.18	0.18	0.36	0.32	0.33	0.11	0.10	0.10	0.34	0.28	0.27	0.57[6.87]
Tunisia	0.09	0.09	0.08	0.09	0.08	0.08	0.06	0.06	0.06	0.11	0.09	0.09	0.26[1.94]

Notes: - The critical value at the 10 percent level is 0.175 and 0.233 at 5 percent level [Sephton (1996) Table 1].

Whether we normalize on the trade or the budget deficit variable, cointegration test results appear mixed. For the trade deficit, the data show that the Harris-Inder test rejects cointegration at the 10 percent level (using eight lags) in Burundi, Rwanda, Sierra Leone, and Togo. In a similar fashion, normalizing on the budget deficit, the test rejects cointegration in Botswana, Burundi, Mauritius, Nigeria, Rwanda, Sierra-Leone and Togo. These results can be interpreted as suggesting that the more appropriate approach is normalizing on the trade deficit. This is so because normalizing on the budget deficit was rejected in all but three cases. It is worth mentioning that these results are generally corroborated by the Shin test results which indicate the presence of no cointegration in Burundi and Rwanda (with trade deficit as the regressand) and Mauritius, Nigeria, Sierra-Leone and Togo (with budget deficit as the regressand). Overall, Shin's test results support normalizing on TD. Also, observe that the results obtained using dynamic least squares (DLS) estimator reveal a positive relationship between the trade deficit and the budget deficit in Africa.

D. Fractional cointegration tests

The empirical results reported above have been obtained using the conventional cointegration approaches, and the evidence has been mixed. The traditional tests have low power when the residuals from a cointegrating equation are mean reverting but not I(0); therefore, greater flexibility can be achieved by using an autoregressive fractionally integrated moving average (ARFIMA) model. An autoregressive fractionally integrated moving average (ARFIMA) process can be postulated as: where d is the long-memory (fractional differencing) parameter that can take on integer and non-integer values, and $\Phi(L)$ and $\Theta(L)$ are the usual autoregressive and moving average polynomials in the lag operator with roots outside the unit circle. It is the fractional differencing parameter in the long-memory models that determines the presence and describes the nature of the long memory.

$$\Phi(L)(1 - L)^d x_t = \Theta(L)\varepsilon_t, \varepsilon_t \sim (0, \sigma^2) \quad (3)$$

An ARFIMA is said to be a stationary process when -0.5 < d < 0.5. This area can be divided into two parts. For -0.5< d < 0 all the autocorrelations are negative and hyperbolic, so the process is considered anti-persistent memory (intermediate memory), whereas for 0 < d < 0.5, the process is called stationary long-memory with positive autocorrelations that decay hyperbolically. If 0.5# d< 1, the process is nonstationary and its mean reversion properties are persistent in nature. Therefore, for 0< d < 1, the process has a long memory (i.e., a fractionally integrated process) and it is mean reverting. In sum, as Cheung and Lai (1993) pointed out, an appealing aspect of the ARFIMA model is its ability to capture a wide variety of low-frequency behavior with a single parameter, d, whose interpretation depends on whether d < 1 or not.[4] Table 3 presents some of the characteristics associated with different estimated values of d.

[4]If d=1 (i.e., I(1) process), it is both covariance nonstationary and not mean-reverting, and an innovation has permanent effect on the series. On the other hand, for a fractionally integrated process with d < 1, an innovation has no permanent effect on the series but its mean-reversion properties are persistent (i.e., the dependence between distant observations is noticeable).

Table 3
Summary of fractional integration values

d value	Mean	Variance	Shock duration	Equilibrium Relationship
d > 1.0	No mean-reversion	Infinite variance	Infinite which effects increases	Cointegration is not valid
d = 1.0	No mean-reversion	Infinite variance	Infinite which effects decreases	Cointegration is not valid
0.5 # d < 1.0	Long run-mean-reversion	Finite variance	Long time	Deviations are mean-reverting but not stationary; Cointegration holds, but very weakly.
0 < d < 0.5	Long run-mean-reversion	Finite variance	Long time	Deviations are mean-reverting and stationary in the long run; Cointegration holds weakly.
d = 0	Long run-mean-reversion	Finite variance	Long time	Deviations follow a stationary and mean-reverting process; Cointegration holds.

Testing for fractional cointegration requires testing for fractional integration in the error-correction term. Cointegration exists when the equilibrium error is mean reverting, that is, d lies inside the limits specified by $0 < d < 1$. In this case, the error term responds slowly to shocks so that deviations from equilibrium are more persistent. To approximate the value of d for the error term, we have opted to use methods suggested by Geweke and Porter-Hudak (1983), henceforth GPH, and Sowell (1992), from amongst the various methods used for implementing cointegration tests based on the estimation of the ARFIMA model.

GPH (1983) suggests the semi-parameteric estimator of d based on the following OLS-based estimating equation:

$$\ln(I(w_j)) = \beta_0 + \beta_1 \ln(4\sin^2(w_j / 2)) + \partial_t ; \quad j = 1,\ldots,n \tag{4}$$

with $\beta_1 = -d$, where $I(w_j)$ is the periodogram of a series at frequency w_j, and $w_j = 2\pi_j / T$ ($j = 1, . . ., T-1$). The number of low-frequency ordinates (n) used in this test is $n=T^{\mu}$, where T is the number of observations. The value of d can be used to test the null hypothesis of a unit root. GPH simulation suggests $\mu = 0.5$ or above. Work by Cheung and Lai (1993) points out that a large number of μ will contaminate the estimate of d while too few will produce imprecise estimates of d. More recent results by Hurvich, Deo and Brosky (1998), Lai (1997) and others find that 0.6# μ # 0.8 are more appropriate values to use in the spectral regression of the GPH (1983) test. In this study we have employed a μ equal to 0.60.

The GPH method of testing for cointegration is a three-step approach. In the first step, the residuals from equation (4) are obtained by means of ordinary least square (OLS). In the second step, to ensure stationarity of the series, the residuals in first differences are generated. In the third step, the GPH test is used to estimate the differencing order of the residuals from the regression of trade deficit on budget deficit and activity measure as well as the residuals from the regression of budget deficit on trade deficit and activity measure. The null d = 1 (no cointegration) is tested against the alternative d < 1 (fractional cointegration). That is, the null hypothesis that trade deficit, budget deficit and activity measure are not cointegrated is tested against the alternative hypothesis that they are cointegrated. The d-estimates from the GPH regression and the t-statistics for the null hypothesis of d=1 are reported in Table 4. Starting with GPH test results for the residuals, it is found that most of the estimated values of d are I(d) with 0 <d < 1.

The exceptions can be found in the results for Burundi and Kenya, where the sign on d is negative, and in the results for South Africa, where the value of d is greater than one. Another interesting feature of the results is that most of the residual series are fractionally integrated since the fractional difference parameter is significantly different from zero, and also a formal hypothesis testing indicates significant evidence of d < 1. On the whole, the results indicate the presence of cointegration and possibly fractional cointegration among trade deficit, budget deficit and activity measure. Therefore, although the series may vary widely, deviations from the cointegrating relationship are mean reverting, so that a shock to the system will eventually die out.

Another appealing aspect of our results is that the GPH test tends to be relatively more powerful than our conventional tests in finding cointegration. For example, for the Harris-Inder test results, we saw that cointegration is rejected in four cases when the regressand is the trade deficit variable. According the GPH test results, cointegration is confirmed for all four cases, namely, Burundi, Rwanda, Sierra-Leone and Togo. In a similar fashion, cointegration which was rejected in seven cases is strongly supported in the GPH test results of d=1. As a cross-check of these results, we report the results obtained using the exact maximum likelihood procedure. The first difference of the residuals was used in the analysis.

Table 4
Cointegration tests using GPH estimator

Name of Country	GPH			
	Residuals normalized on trade deficit		Residuals normalized on budget deficit	
	p=0.6		*p=0.6*	
	H0: d = 0	*H0: d = 1*	*H0: d = 0*	*H0: d = 1*
Botswana	*0.338[1.69, 1.66]*	*0.338[3.13, 3.24]*	*0.067[0.37, 0.33]*	*0.067[5.11, 4.57]*
Burundi	*0.525[4.89, 2.36]*	*0.525[4.42, 2.14]*	*-0.083[0.37, 0.37]*	*-0.083[4.83, 4.87]*
Kenya	*-0.020[0.11, 0.10]*	*-0.020[5.72, 5.00]*	*0.238[1.49, 1.17]*	*0.238[4.77, 3.74]*
Mauritius	*0.428[2.61, 2.01]*	*0.428[3.49, 2.69]*	*0.950[8.90, 4.47]*	*0.950[0.47, 0.24]*
Nigeria	*0.335[1.40, 1.71]*	*0.335[2.79, 3.39]*	*0.676[2.80, 3.44]*	*0.676[1.34, 1.65]*
Rwanda	*0.220[1.10, 1.04]*	*0.220[3.91, 3.67]*	*0.376[1.42, 1.77]*	*0.376[2.36, 2.94]*
Sierra- Leone	*0.586[4.33, 2.99]*	*0.586[3.06, 2.11]*	*0.394[1.76, 2.01]*	*0.394[2.70, 3.09]*
South- Africa	*1.071[3.31, 5.46]*	*1.071[0.22, 0.36]*	*0.622[3.09, 3.17]*	*0.622[1.88, 1.93]*
Togo	*0.358[1.28, 1.68]*	*0.358[2.30, 3.02]*	*0.225[1.63, 1.06]*	*0.225[5.61, 3.65]*
Tunisia	*0.078[0.70, 0.38]*	*0.078[8.23, 4.52]*	*0.346[1.83, 1.70]*	*0.346[3.45, 3.20]*

Notes: -The critical values for the GPH are 2.10 at the 5 percent level and 1.59 at the 10 percent level (Anderson and Lyhagen (1997)). The first value in square bracket is t-value based on the empirical error variance (OLS) and the second value is Z based on theoretical error variance. For example, using Burundi's data, the OLS standard error is 0.1074 (0.525/4.89) and the theoretical standard error is 0.2224 (0.525/2.36). To test H_0: d = 1, we obtain in absolute term 4.42 [i.e., (0.525-1)/0.1074] and 2.14[(0.525-1/0.2224)], respectively.

Table 5 reports the parameter estimates for various models obtained using the exact maximum likelihood procedure. We experimented with numerous alternative ARFIMA (p,d*,q) with p = 0, 1, 2 and q =0, 1, 2 . The Akaike Information Criterion (AIC) procedure was used for selecting reasonable models.

For all selected models, standard t-tests were computed and have been reported in Table 5. Normal distribution is used for critical values in the t-test. One can be added to the estimates for d* in order to obtain estimates for d. Looking at the data in Table 5, it can easily be seen that the results are qualitatively the same as those obtained from using the GPH estimator. Moreover, the residual series are fractionally integrated as the null hypothesis of no fractional cointegration (d* = 0) is rejected in most cases.

An important issue relates to the contribution of activity measure (ZT) to the fractional cointegration property of the estimated models. To measure that, the models were estimated after dropping the activity variable. The results are in the appendix, and they convincingly show that the inclusion of the activity variable did not influence the reported fractional cointegration test results.

Table 5

Fractional cointegration tests using Maximum Likelihood Estimator

(With Trade Deficit, Budget Deficit and Activity)

Name of Country	Maximum Likelihood ARFIMA (p,d,q)	d*	T-Value	Maximum Likelihood ARFIMA (p,d,q)	d*	T-Value
	Residuals normalized on trade deficit			Residuals normalized on budget deficit		
Botswana	(0,d,1)	-0.584	[5.25]	(0,d,1)	-0.935	[3.43]
Burundi	(0,d,1)	-0.335	[2.56]	(0,d,0)	-0.956	1[0.10]
Kenya	(0,d,1)	-0.481	[4.39]	(0,d,1)	-0.886	[4.49]
Mauritius	(0,d,1)	-0.459	[3.95]	(0,d,1)	-0.443	[4.12]
Nigeria	(0,d,1)	-0.583	[6.15]	(0,d,0)	-0.628	[8.75]
Rwanda	(0,d,1)	-0.835	[6.77]	(0,d,1)	-0.537	[4.42]
Sierra Leone	(0,d,1)	-0.441	[4.65]	(0,d,1)	-0.337	[2.82]
South- Africa	(0,d,1)	-0.155	[0.99]	(0,d,1)	-0.562	[3.48]
Togo	(0,d,1)	-0.563	[4.62]	(0,d,1)	-0.466	[3.62]
Tunisia	(0,d,1)	-0.517	[4.43]	(0,d,1)	-0.570	[4.93]

Notes: - ARFIMA (p,d,q) models are estimated for each residual data series using p q = 0, 1, 2. ARFIMA modelling was done using PcGive software.

E. Granger causality tests

In this section, we examine the issue of causality. Engle and Granger (1987) demonstrates the general duality between cointegration and the VECM. Further, they show that cointegration implies that standard Granger causality tests are misspecified and that use should be made of error-correction models instead. These make the VECM an ideal tool to examine Granger causality among budget deficit, trade deficit and activity, as well as the speed of convergence of the relevant variables to their equilibrium. By the Granger representation theorem (GRT) and by focusing on TD and

BD, the relevant error-correction models are as follows:

$$\Delta TD_t = a_0 + \lambda_1 \varepsilon_{t-1} + \sum_{i=1}^{n} a_{1i} \Delta TD_{t-i} + \sum_{i=1}^{n} a_{2i} \Delta BD_{t-i} + \sum_{i=1}^{n} a_{3i} \Delta y_{t-i} + \nu_{1t} \qquad (5)$$

$$\Delta BD_t = b_0 + \lambda_2 \varepsilon_{t-1} + \sum_{i=1}^{n} b_{1i} \Delta TD_{t-i} + \sum_{i=1}^{n} b_{2i} \Delta BD_{t-i} + \sum_{i=1}^{n} b_{3i} \Delta y_{t-i} + \nu_{2t} \qquad (6)$$

where $\lambda 1$ and λ_2 are coefficients for the error-correction terms in equations (5) and (6). These coefficients are expected to capture the adjustments of ΔTD_t and ΔBDt towards long-run equilibrium. In our case, equation (5) is used to test causation from budget deficits to trade deficits – (that is, budget deficits do not Granger-cause trade deficits if all a_{2i} = 0 and /or $\lambda 1 = 0$), while equation (6) is used to test causality from trade deficits to budget deficits - trade deficits do not Granger-cause budget deficits if all b_{1i} =0 and/or $\lambda_2 = 0$. Since we found evidence of cointegration, there must be either unidirectional or bidirectional Granger causality, because at least one of the error correction terms should be significantly different from zero by the definition of cointegration.

The VECM approach, besides showing the direction of Granger-causality among the variables, enables one to distinguish between 'short-run' and 'long-run' Granger-causality. The former is generally referred to as the channel 1 source of causation and can be evaluated by testing whether the estimated coefficients on lagged values of ΔBD_t [all a_{2i} =0] in equation (5) or lagged values of ΔTD_t [all b_{1i} = 0] in equation (6) are jointly statistically significant. This can be done using a standard Wald test. For convenience, we interpret this 'short-run' Granger-causality as weak causality. On the other hand, 'long-run' Granger causality is generally referred to as the channel 2 source of causation and can be evaluated by testing whether the coefficient of the error-correction term in each equation [that is, $\lambda_1 = 0$; $\lambda_2 = 0$] is statistically different from zero by a t-test. For consistence, we have used a Wald test.

When testing for Granger causality, it is desirable to examine whether both channel 1 and channel 2 sources of causation are jointly significant. This is so because the joint test will indicate which variable(s) bear the burden of short-run adjustment to reestablish long-run equilibrium, following a shock to the system (Asafu-Adjaye, 2000). This is referred to as a 'strong causal relation' and can be examined by testing the joint null hypothesis that all a_{2i} = 0 and $\lambda_1 = 0$ in equation (5) or that all b_{1i} = 0 and $\lambda_2 = 0$ in equation (6).

The empirical results of causality through these channels are shown in Table 6. We report three causality tests relating to zero restriction of relevant variables in the VECM where the null hypothesis is that there is no Granger causality against the alternative that there is Granger causality.

In Table 6 beginning with the short-run Granger causality, the Wald statistics suggest that budget deficits Granger-cause trade deficits in four countries (Botswana, Kenya, Nigeria and South Africa) at the 10 percent level or better. Similarly, there is evidence of short-run Granger causality running from trade deficits to budget deficits in one country (Rwanda). For Togo short-run bidirectional Granger causality was observed, while no evidence of short-run causal relation between trade deficits and budget deficits was found in Tunisia.

Table 6
Granger test of causality: Wald test statistics
African Countries

		$H_0 : \Delta TD_t \leftarrow \Delta BD_t$			$H_0 : \Delta TD_t \rightarrow \Delta BD_t$		
		Source of causation (independent variable)					
		Short run	Long run		Short run	Long run	
Name of Country	**lags**	**ΔBD_t lags**	ΔBD_t lags and $\in_{t-1}$	$\in_{t-1}$	**ΔTD_t lags**	**ΔTD_t lags and** $\in_{t-1}$	$\in_{t-1}$
Botswana	5	9.48(0.05)	13.58(0.02)	9.09(0.00)	1.49(0.83)	18.20(0.00)	6.64(0.01)
Burundi	3	1.09(0.58)	3.88(0.27)	0.34(0.56)	3.27(0.20)	22.45(0.00)	20.07(0.00)
Kenya	5	7.95(0.09)	7.66(0.18)	5.76(0.02)	2.11(0.72)	11.94(0.04)	9.17(0.00)
Mauritius	2	0.98(0.32)	25.80(0.00)	25.38(0.00)	0.03(0.87)	7.77(0.02)	6.68(0.01)
Nigeria	3	12.69(0.00)	22.44(0.00)	17.43(0.00)	3.52(0.17)	5.92(0.12)	4.13(0.04)
Rwanda	5	1.84(0.77)	30.20(0.00)	25.96(0.00)	27.23(0.00)	49.43(0.00)	6.25(0.01)
Sierra Leone	6	6.29(0.28)	12.28(0.06)	7.00(0.01)	5.78(0.33)	17.66(0.01)	11.92(0.00)
South Africa	4	6.23(0.10)	7.68(0.10)	0.33(0.57)	5.67(0.13)	17.76(0.00)	10.75(0.00)
Togo	5	54.95(0.00)	56.67(0.00)	3.03(0.08)	12.43(0.01)	118.19(0.00)	84.67(0.00)
Tunisia	7	2.83(0.83)	4.22(0.75)	3.23(0.07)	6.15(0.41)	10.13(0.18)	6.81(0.01)

Note: Lags are equal to (k-1) for the ECM.

With respect to the long-run Granger causality, there is strong evidence of bidirectional causality (feedback) between trade deficits and budget deficits in eight cases (Botswana, Kenya, Mauritius, Nigeria, Rwanda, Sierra Leone, Togo, and Tunisia). The results for two cases (Burundi and South Africa) confirm unidirectional Granger-causality running from trade deficits to budget deficit.

Concerning the issue of burden of adjustment towards the long-run equilibrium in response to a short-run deviation, the most striking result is that both BD and TD are important for most of the cases (Botswana, Mauritius, Rwanda, Sierra Leone, South Africa and Togo) in our sample. This verdict, in a multi-country framework is corroborated by the work of Darrat (1988), Kouassi et al. (2004) and several others. We also found that the burden of adjustment depended on TD in two cases (Burundi and Kenya), whereas in Nigeria the burden of adjustment appears to depend on only BD.

In sum, for a majority of the cases studied, long-run bidirectional Granger causality was dictated between TD and BD; however, in the case of short-run Granger causality, the results revealed one case (Togo) of bidirectional causality. It is one-way causality from BD to TD (Botswana, Kenya, Nigeria and South Africa) and as well as no causality (Mauritius, Sierra Leone and Tunisia) between the budget and trade deficits that are tended to characterize the short-run dynamics. Short-run unidirectional Granger causality from TD to BD was obtained for Rwanda and bidirectional Granger causality was obtained for Togo.

F. Speed of adjustment and mean time-lags

Studies of the relationship between budget deficits and trade deficits are notably silent on issues regarding the adjustment mechanism of their convergence process. This section provides information on the convergence speeds of the trade deficit and the budget deficit. Using vector error-correction (VEC) analysis, we estimate the speeds at which the individual variables revert to their long-run values as well as the mean time of the response.

The speed of adjustment is represented by the absolute value of the error-correction term, which can be interpreted as the response of the dependent variable in each period to the departures from equilibrium, or for example, as the change in the budget deficit variable per quarter that is attributed to the disequilibrium between the actual and the equilibrium levels. As the data in Table 7 show, there is considerable inter-country variation in the adjustment speed to the last period's disequilibrium. When the TD is the regressand, the coefficient of the error-correction term ranges (in absolute term) from a low of 0.052 for Burundi to a high of 0.407 in Tunisia, whereas for BD as the regressand, the range is from a low of 0.117 for Mauritius to a high of 0.954 in Botswana. For example, in Mauritius only 11.7 percent of adjustment occurs in a quarter, whereas the figure is 95.4 percent for Botswana.

Since the relative size of the error-correction coefficients in the VECM governs to a large extent the adjustment of the individual variables to the equilibrium, this allows us to confirm that the budget deficit is the primary adjuster to disequilibria by presenting evidence that the size of the error-correction coefficient in the budget deficit equations of the VECM are of an order of magnitude larger in absolute value than those in the trade deficit equation.

The median (half-life) time lag gives the time for 50 percent of the adjustment of regressand to changes in the regressors to occur. For the budget deficit equation, the median time lag ranges from a low of 0.22 quarters in Botswana to a high of 5.60 quarters (i.e., 1.4 years) for Mauritius. The average of the median (half-life) adjustment is 1.88 quarters for our sample. These estimates are fairly short relative to those of the trade deficit equation, where the median time lag ranges from a low of 1.33 in Tunisia to a high of 13. 01 quarters (3.25 years) in Burundi, with the average half-life adjustment being 3.74 quarters.

Table 7
Speed of Adjustment Estimates

Name of Country	Trade deficit as dependent				Budget deficit as dependent			
	εt-1	Overall median time lag	Mean time lag		εt-1	Overall median time lag	Mean time lag	
			BD	*y*			*TD*	*y*
Botswana	*-0.165 (0.077)*	*3.84*	*4.99*	*6.08*	*-0.954 (0.349)*	*0.22*	*1.44*	*1.82*
Burundi	*-0.052 (0.053)*	*13.01*	*15.04*	*18.43*	*-0.950 (0.384)*	*0.23*	*0.78*	*0.40*
Kenya	*-0.288 (0.079)*	*2.04*	*1.94*	*1.98*	*-0.939 (0.250)*	*0.25*	*1.59*	*1.28*
Mauritius	*-0.234 (0.105)*	*2.62*	*2.68*	*3.58*	*-0.117 (0.099)*	*5.60*	*15.92*	*18.67*
Nigeria	*-0.280 (0.106)*	*2.11*	*5.04*	*3.72*	*-0.289 (0.122)*	*2.03*	*3.88*	*4.15*
Rwanda	*-0.239 (0.162)*	*2.53*	*13.39*	*12.56*	*-0.227 (0.136)*	*2.70*	*5.36*	*6.38*
Sierra Leone	*-0.136 (0.058)*	*4.75*	*4.39*	*7.41*	*-0.157 (0.063)*	*4.06*	*4.14*	*4.12*
South Africa	*-0.214 (0.081)*	*2.88*	*4.44*	*2.59*	*-0.537 (0.175)*	*0.90*	*4.75*	*5.54*
Togo	*-0.265 (0.097)*	*2.25*	*2.60*	*2.45*	*-0.464 (0.081)*	*1.11*	*2.28*	*0.99*
Tunisia	*-0.407 (0.120)*	*1.33*	*0.72*	*0.23*	*-0.330 (0.114)*	*1.73*	*1.22*	*1.44*
Average		*3.74*	*5.52*	*5.90*		*1.88*	*4.14*	*4.48*

The mean time lag gives the mean (average) length of time for full-adjustment (i.e., the average length of time it takes for a (unit) change in the regressors to be transferred to the regressand. For most of the cases examined, the mean time lag for the response of the budget deficit to changes in the trade deficit is shorter than the response of the trade deficit to changes in the budget deficit. Also, for the former, the average mean time lag is 4.23 quarters, whereas for the latter, it is 5.52 quarters. In sum, the adjustment is made by the budget deficit. In other words, it is the budget deficit that adjusts its level for the two variables to share the same relationship over time.

IV. SUMMARY AND CONCLUSION

In this paper, the relationship between the trade deficit and the budget deficit for ten African countries is investigated using popular time series methodologies. When we used conventional contegration techniques and quarterly data (1973:2 through 2005:4), we obtained mixed evidence concerning cointegration. However, when we relaxed the condition in the conventional cointegration analyses that the residual series must be I(0), we were able to obtain results supportive of fractional cointegration – this generalized form of cointegration captures slow or subtle mean-reverting dynamics -- in almost all cases. Also, in terms of our long-run analyses we gather that economic agents in these African economies are found not to follow Ricardian Equivalence behavior since the two deficits appear to be related.

To determine whether there is a causal relationship between the trade deficit and the budget deficit, the vector error-correction model, where lagged error-correction term is I(d) or fractional, was estimated. In contrast to the results in previous studies, bidirectional long-run causality between the budget deficit and the trade deficit receives strong empirical support in the majority of cases. Furthermore, the results show that the source of causation in the long-run is the error-correction terms in both direction.

The results from Granger causality tests on vector-correction models suggest that, in comparison to the long-run relation, the short-run dynamics reflect largely one-way causality running from the trade deficit to the budget deficit as well as no causality between the deficits. Bidirectional short-run Granger causality was observed in only one case.

On the basis of the relative magnitudes of the budget deficit error-correction coefficients versus the trade deficit error-correction coefficients, the empirical results reveal that the budget deficit is the primary variable that changes in order to restore equilibrium when the system has been subjected to some shock. Also, both the mean and the median time lags for the response of the budget deficit to changes in the trade deficits are shorter than those of the response of the trade deficit to changes in the budget deficit.

All in all, the results of this study have been good in the sense that common equations were able to perform well for a diverse sample of countries. The results shed light on the nature of the relationship between the trade deficit and the budget deficit in these developing countries and suggest that balancing the budget cannot be expected to reduce the trade deficit; however, what is needed are both fiscal and monetary policies. In addition, the results suggest that government policy actions aimed at reducing trade and budget deficits can generate at best only uncertain results if the initial focus is not placed on policies that attempt to reduce the trade deficit.

REFERENCES

AlKswani, M. A. (2005). The Twin Deficits Phenomenon in Petroleum Economy: Some Evidence From Saudi Arabia. Retrieved in January 26, 2005, from http://www.erf.org.eg/html/finance.pdf.

Anoruo, E. and Ramchander, S. (1998). Current Account and Fiscal Deficits: Evidence from Five Developing Economics of Asia. *Journal of Asian Economics*, Vol. 9, No. 3, pp. 487-501.

Anderson, J. and Lyhagen, J. (1997). A Note on the Power of the GPH Test for Cointegration. Research Report #7 Department of Statistics, Uppsala University.

Arize, A. C.; Osang, T. and Slottje, D. J. (2000). Exchange-Rate Volatility and Foreign Trade: Evidence From Thirteen LDCs. *Journal of Business & Economic Statistics*, Vol. 18, pp. 10-17.

Arize, A. C. and Ndubizu, G. (1992). Error-Correction Representation and Import Demand Function with Implications for International Finance & Accounting. *Review of Quantitative Finance and Accounting*, Vol. 2 (Dec.), pp. 359-376.

Arize, A. C. and Shwiff, S. S. (1998). The Appropriate Exchange-Rate Variable in Money Demand of Twenty-Five Countries: An Empirical Investigation. *North American Journal of Economics and Finance*, Vol. 9, pp. 169-185.

Cheung, Y. W and Kon, S. L. (1993). A Fractional Cointegration Analysis of Purchasing Power Parity. *Journal of Business and Economic Statistics*, Vol. 11, pp. 103-112.

Darrat, A. F. (1988). Have Large Budget Deficits Caused Rising Trade Deficits. *Southern Economic Journal*, Vol. 54, No. 4, pp. 879-887.

Engle, R. F. and Granger, C. W. J. (1987). Cointegration and Error Correction: Representation, Estimation and Testing. *Econometrica*, Vol. 55, pp. 251-276.

Geweke, J. and Porter-Hudak, S. (1983). The Estimation and Application of Long Memory Time Series Models. *Journal of Time Series Analysis*, Vol. 4, pp. 221-238.

Harris, David, and Brett Inder (1994). A Test of the Null Hypothesis of Cointegration. Nonstationary Time Series Analysis and Cointegration, edited by Colin Hargreaves, Oxford: Oxford University Press.

Hurvich, C. M.; Deo, R. S. and Brodsky, J. (1998). The Mean Squared Error of Geweke and Porter-Hudak's Estimator of the Memory Parameter of A Long Memory Time Series. *Journal of Time Series Analysis*, Vol. 19, pp. 19-46.

Kearney, Colm and Monadjemi Mehdi (1990). Fiscal Policy and Current Account Performance: International Evidence on the Twin Deficits. *Journal of Macroeconomics*, Vol 12, pp. 197-220.

Khalid, A. M. and Guan, T. W.(1999). Causality Tests of Budget and Current Account Deficits: Cross-Country Comparisons. *Empirical Economics*, Vol. 24, pp. 389-402.

Kim. C. and Donggeum, K. (2006). Does Korea have Twin Deficits? *Applied Economic Letters*, Vol. 13, pp. 675-680.

Kouassi, Eugene, Mbodja Mougoue and Kymn Kern O (2004). Causality Tests of the Relationship Between The Twin Deficits. *Empirical Economics*, Vol. 29, pp. 503-525.
Lai, K. S. (1997). Long-term Persistence in the Real Interest Rate: Some Evidence of A Fractional Unit Root. *International Journal of Finance and Economics*, Vol. 2, pp. 225-235.
Lutkepohl, H. (1983). Non-Causality due to Omitted Variables. *Journal of Econometrics*, Vol. 19, pp. 367-78.
Lutkepohl, H. (1993). Introduction to Multiple Time Series Analysis. 2nd edition, Berlin: Springer-Verlag.
Sephton, P.S. (1996). Extended Critical Values For A Simple Test of Cointegration. *Applied Economics Letters*, Vol. 3, pp. 55-157.
Shin, Y. (1994). A Residual-Based Test of the Null of Cointegration Against the Alternative of Cointegration. *Econometric Theory*, Vol. 10, No. 1, pp. 91-115.
Sowell, F. B. (1992). Maxiimum Likelihood Estimation of Stationary Univariate Fractionally Integrated Time Series Models. *Journal of Econometrics*, Vol. 53, pp. 165-188.
Summers, L. H. (1988). Tax Policy and International Competitiveness,'in J. Frenkel (eds.) International Aspects of Fiscal Policies, Chicago: Chicago UP, 349-375.

APPENDIX A

Fractional cointegration tests using Maximum Likelihood Estimator
(With Trade Deficit and Budget Deficit)

Countries	Maximum Likelihood ARFIMA (p,d,q)	d*	T-Value	Maximum Likelihood ARFIMA (p,d,q)	d*	T-Value
	Residuals normalized on trade deficit			Residuals normalized on budget deficit		
Botswana	(0,d,1)	-0.539	[5.11]	(0,d,1)	-0.297	[1.40]
Burundi	(0,d,1)	-0.290	[2.69]	(0,d,0)	-0.846	[9.41]
Kenya	(0,d,1)	-0.462	[4.26]	(0,d,1)	-0.983	[5.06]
Mauritius	(0,d,1)	-0.438	[3.71]	(0,d,1)	-0.472	[4.07]
Nigeria	(0,d,1)	-0.558	[6.20]	(0,d,1)	0.384	[5.21]
Rwanda	(0,d,1)	-0.715	[5.98]	(0,d,1)	-0.485	[4.05]
Sierra Leone	(0,d,1)	-0.419	[4.62]	(0,d,1)	-0.307	[2.58]
South- Africa	(0,d,1)	-0.156	[1.05]	(0,d,1)	-0.493	[3.46]
Togo	(0,d,1)	-0.544	[4.88]	(0,d,1)	-0.465	[3.84]
Tunisia	(0,d,1)	-0.483	[4.26]	(0,d,1)	-0.505	[4.71]