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Review

Cointegrating and causal relationship between financial development and economic growth in ECOWAS countries

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The aim of this paper is to re-examine the cointegrating and causal relationship between financial development and economic growth in the ECOWAS countries. To this end, we use the Pesaran et al. (2001) approach to cointegration and the procedure for non - causality test of Toda and Yamamoto (1995). Data are from the World Bank (2007) and cover the period 1960 - 2005. We show that there is a positive long - run relationship between financial development and economic growth in five countries, namely, Cape Verde, Cote d'Ivoire, Ghana, Guinea and Liberia. In addition, we show that financial development 'leads' economic growth in Ghana, Liberia and Mali while growth causes finance in Cote d'Ivoire, and a bidirectional causality in Cape Verde and Sierra Leone. The policy implication is that Cape Verde, Ghana, Liberia, Mali and Sierra Leone should give policy priority to financial reform while Cote d'Ivoire should promote economic growth.

Key words: Cointegration, financial development, granger causality, growth.

INTRODUCTION

Every economy requires a sophisticated and efficient financial system to prosper since a healthy financial system is integral to the sound fundamentals of an economy. A more efficient financial system provides better financial services, and this enables an economy to increase its gross domestic product (GDP) growth rate. Hence, in the last decades, many developing countries, particularly West African countries, have adopted development strategies that prioritize the modernization of their financial systems. Since the end of the 1980s, the ECOWAS countries have implemented reforms policies in their financial systems within the context of structural adjustment proposed by the Bretton Woods institutions. These reforms ought to foster financial development through the reduction of governmental intervention in national financial sectors or the privatization of banks. Such policies have been expected to promote growth through, among others, higher mobilization of savings or a rise in domestic and foreign investments (Gries et al., 2009). However, the effectiveness of such policies requires a convenient causal relationship between financial and real sectors.

JEL Classification: C32, E44.

The relationship between financial development and economic growth has received considerable attention in the theoretical and empirical literature. However, economists disagree sharply about the role of financial sector in economic growth. The debate has traditionally revolved around three issues. The first view suggests that the increase in the demand for financial services resulting from economic growth is the major driving force behind the development of the financial sector. This mechanism is stressed in the work of Robinson (1952). According to this strand of literature, financial development follows economic growth or 'where enterprise leads finance follows'. In other words, as the real side of the economy expands, its demand for financial services increases, leading to the growth of these services. Empirical support for this view can also be found in some recent studies (Demetriades and Hussein, 1996). The second view, proposed by Schumpeter (1912), Goldsmith (1969), Hicks (1969), McKinnon (1973), Gurley and Shaw (1955), requires a convenient causal relationship between

Miller (1998), emphasizes a proactive role for financial services in promoting economic growth. In this view, financial development has a positive effect on economic growth. In other words, financial intermediation contributes to economic growth through two main channels: by raising the efficiency of capital accumulation and in turn the marginal productivity of capital and by raising the savings rate and thus the investment rate. A last view provided by Lucas (1988) dismisses finance as an 'over - stressed' determinant of economic growth or in other words financial development and economic growth are not causally related. All these points of view are recently reviewed by Levine (2005).

Recent empirical analyses of the influence on long - run economic growth of financial development include, for example, Levine (1999), Aghion et al. (2005), Levine et al. (2000), Roubini and Sala-i-Martin (1992), King and Levine (1993). These studies used cross - section analysis to link measures of financial development with economic growth. Cross - country growth regressions do not capture the dynamics of the relationship between financial development and economic growth. In addition, a significant coefficient of financial development in growth regressions does not necessarily imply causality running from finance to growth or vice versa. Such improper assessments of causal relationships in a static cross - section setting have led researchers to seek more dynamic time series analyses to unravel whether financial development causes economic growth or vice versa. Moreover, many other studies have highlighted the inappropriateness of cross - sectional analysis. Hence, time series studies of a selection of countries by Abu-Bader and Abu-Qarn (2008), Al-Yousif (2002) or Demetriades and Hussein (1997) have shown that the pattern of causality differs significantly among countries that strengthen the lead of country - specific studies.

The aim of this paper is to study the cointegrating and causal relationship between financial development and economic growth in the Economic Community of West African States ([1] ECOWAS is composed of Benin, Burkina Faso, Cape Verde, Cote d'Ivoire, Gambia, Ghana, Guinea, Guinea-Bissau, Liberia, Mali, Niger, Nigeria, Senegal, Sierra Leone and Togo). (ECOWAS). This is an important concern because it assists in an evaluation of the extent to which the development of financial sector has spurred economic growth in the ECOWAS area. Further, it gives some guidance as to whether financial sector development is a necessary and sufficient condition for a higher growth rates in developing countries. To this end, we follow the Pesaran et al. (2001) approach to cointegration and the Toda and Yamamoto's (1995) procedure to test for the non - causality between the variables of interest. The Pesara et al. (2001) approach has at least two major advantages over the traditional approaches (Engle and Granger, Johansen) used by a wide range of studies. The first advantage is that it is applicable irrespective of whether the underlying regressors are purely stationary, purely integrated 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. Furthermore, Toda and Yamamoto (1995) propose an interesting yet simple procedure requiring the estimation of an augmented vector autoregressive (VAR) which guarantees the asymptotic distribution of the Wald statistic, since the testing procedure is robust to the integration and cointegration properties of the process. Data are from the 2007 world development indicators of the World Bank (2007) and cover the period 1960 - 2005.

Following standard practice, we use real gross domestic product (GDP) per capita as our measure for economic growth. In the line of recent works, the ratio of credit to private sector to GDP has been used as measure of financial development. We build unrestricted error correction model including a measure of economic growth and a financial development indicator and test for the null of no long - run link between these two variables. In addition, we construct bivariate levels vector autoregressive model and test for the non - causality from financial development to economic growth, and vice versa. Pesaran et al. (2001) cointegration test results show the existence of a long run relationship between financial development and economic growth in Cape Verde, Cote d'Ivoire, Ghana, Guinea and Liberia. Furthermore, following Toda and Yamamoto (1995), there is a bidirectional causality in Cape Verde and Sierra Leone; financial development significantly causes economic growth in Ghana, Guinea, Liberia and Mali while growth causes finance in the case of Cote d'Ivoire. The remainder of this paper is organized as follows. Section 1 highlights the econometric framework. In the Section 2, we present the main results of this study. Section 3 provides some policy implication. We finish by the conclusion.

THE ECONOMETRIC FRAMEWORK

This section highlights the econometric framework used to study cointegration and causality between financial development and growth. We use the Pesaran et al. (2001) cointegration approach and the Toda and Yamamoto (1995) causality testing procedure.

The cointegration analysis

Econometric literature proposes different methodological alternatives to empirically analyse the long - run relationships and dynamics 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 (1988) and Johansen and Juselius (1990). All these methods require that the variables under investigation are integrated of order one. This inevitably involves a step of stationarity 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; Harris, 1995). Due to these problems, this study makes use of a newly developed approach to cointegration that has become popular in recent years.

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 used by a wide range of studies (Masih and Masih 2000; Narayan and Peng, 2007). 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. This test is particularly appropriate for small samples in which the order of integration is not known or may not be necessarily the same for all variables of interest. To search for possible long run relationships amongst the variables, namely gross domestic product per capita (denoted by *GDPC*) and financial development (credit to private sector as a percentage of GDP, denoted by *CRE*), we employ the bounds testing approach to cointegration suggested by Pesaran et al. (2001). This involves estimating the following unrestricted error correction model (UECM):

$$\Delta Y_t = \alpha_0 + \alpha_1 Y_{t-1} + \alpha_2 F_{t-1} + \sum_{i=1}^{m} \beta_i \Delta Y_{t-i} + \sum_{i=0}^{m} \gamma_i \Delta F_{t-i} + \varepsilon_t \quad (1)$$

where $Y = \ln(GDPC)$, $F = \ln(CRE)$, α_i s $(i = 0, 1, 2)$, β_i s $(i = 1, 2, ..., m)$, and γ_i s $(i = 0, 1, 2, ..., m)$, are the parameters of the model. The structural lags m are determined by using minimum Akaike information criteria (AIC) and Schwarz Bayesian information criteria (SC). To depict the presence of cointegration the estimated coefficients of lagged level variables are restricted equal to zero. Thus, the null hypothesis of no cointegration between finance and economic growth according to equation (1) can be expressed as:

$$H_0 : \alpha_1 = \alpha_2 = 0 \quad (2)$$

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 (30 to 46 observations) we calculate critical values specific to our sample sizes. To this end, we use a GAUSS code to generate the original set of critical values. These critical values are computed using stochastic simulations for different sample sizes *T* = 30, 32, 34, 36, 39, 40, 44, 46, 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 (a model with an intercept and an unrestricted trend is referred to as Case V), and is expressed as:

$$\Delta y_t = \delta_0 + \delta_1 y_{t-1} + \delta_2 x_{t-1} + \varepsilon_t \quad (3)$$

Here $t = 1, 2, ..., T$, $z_{t-1} = \left(y_{t-1}, x_{t-1}\right)'$, $w_t = 1$. The variables y_t and x_t are generated from $y_t = y_{t-1} + \varepsilon_{1t}$, and $x_t = Px_{t-1} + \varepsilon_{2t}$, with $y_0 = 0$, $x_0 = 0$ and $\varepsilon_t = \left(\varepsilon_{1t}, \varepsilon_{2t}\right)'$ is drawn as two independent standard normal variables. If x_t is purely I(1), that is, integrated of order one, $P = 1$. 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 financial development and GDP per capita 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 (Narayan, 2005).

If a cointegration relationship is observed between the series, Bardsen's (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 finance and growth is described by:

$$Y_t = \vartheta_0 + \vartheta_1 F_t + \mu_t \tag{4}$$

where $Y = \ln(GDPC)$, $F = \ln(CRE)$, $\vartheta_0 = -\frac{\alpha_0}{\alpha_1}$ and $\vartheta_1 = -\frac{\alpha_2}{\alpha_1}$, and μ_t is a stationary process with zero mean.

The Toda and Yamamoto 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 inappropiate 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 GDP per capita and the credit to private sector ratio, following Yamada (1998), and examine the non-causality between these variables:

$$Y_t = \varphi_0 + \sum_{i=1}^{p} \psi_i Y_{t-i} + \sum_{i=p+1}^{p+d_{\max}} \psi_i Y_{t-i} + \sum_{i=1}^{p} \varphi_i F_{t-i} + \sum_{i=p+1}^{p+d_{\max}} \varphi_i F_{t-i} + \nu_{1t} \tag{5}$$

$$F_t = \chi_0 + \sum_{i=1}^{p} \eta_i F_{t-i} + \sum_{i=p+1}^{p+d_{\max}} \eta_i F_{t-i} + \sum_{i=1}^{p} \chi_i Y_{t-i} + \sum_{i=p+1}^{p+d_{\max}} \chi_i Y_{t-i} + \nu_{2t} \tag{6}$$

Where $Y = \ln(GDPC)$, $F = \ln(CRE)$, φ_i s, ψ_i s, η_i s, and χ_i s are the parameters of the model; $d_{\max}$ is the maximum order of integration suspected to occur in the system; $\nu_{1t} \sim N(0, \Sigma_{\nu 1})$ and $\nu_{2t} \sim N(0, \Sigma_{\nu 2})$ 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 $\ln(CRE)$ to $\ln(GDPC)$ can be expressed as:

$$H_0 : \varphi_i = 0, \ \forall \ i = 1, 2, ..., p \tag{7}$$

where the φ_i are the coefficients of the lagged values of ln(*CRE*) in the growth equation. Let $\varphi = vec(\varphi_1, \varphi_2, ..., \varphi_p)$ 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:

$$W = T\left(\hat{\varphi}' R' \left(R \hat{\Sigma}_\nu R' \right)^{-1} R \hat{\varphi} \right), \tag{8}$$

where $\hat{\varphi}$ is the ordinary least squares estimate for the coefficient φ and $\hat{\Sigma}_\nu$ is a consistent estimate for the asymptotic covariance matrix of $\sqrt{T}(\hat{\varphi} - \varphi)$. The test statistic is asymptotically distributed as a χ^2 with p degrees of freedom.

Two steps are involved with implementing the procedure. The first step includes determination of the lag length (p) and 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 Zivot and Andrews (1992) test to determine the maximum order of integration.

The Zivot and Andrews (1992) unit root test

A break in the deterministic trend affects the outcome of unit root tests. Several studies have found that the conventional unit root tests fail to reject the unit root hypothesis for series that are actually trend stationary with a structural break. Perron (1989) showed that a Dickey and Fuller (1979) type test for unit root is not consistent if the alternative is that of a stationary noise component with a break in the slope of the deterministic trend. His main point is that the existence of exogenous shock which has a permanent effect will lead to a non-rejection of the unit root hypothesis even though it is true. Perron (1989, 1990) proposed alternative unit root tests which allow the possibility of a break under the null and alternative hypotheses. They have less power than the Dickey-Fuller test when there is no break but they are consistent when there is a break or not. Furthermore, they are invariant to the break and parameter and thus their performance does not depend on the magnitude of the break. However, the most controversial assumption is that its timing is known a priori (Christiano, 1992). The use of an incorrect break date in Perron's (1990) tests causes size distortions and power loss, though this effect disappears asymptotically (Kim and Perron, 2009). The work by Zivot and Andrews (1992) provides methods that treat the occurrence of the break date as unknown. To test for a unit root against the alternative of trend stationary process with a structural break, the following regressions are used:

$$Model\ A:\ y_t = \mu + \theta DU_t(\tau_b) + \beta t + \alpha y_{t-1} + \sum_{i=1}^{k} \varphi_i \Delta y_{t-i} + e_t \qquad (9)$$

$$Model\ B:\ y_t = \mu + \beta t + \gamma DT_t(\tau_b) + \alpha y_{t-1} + \sum_{i=1}^{k} \varphi_i \Delta y_{t-i} + e_t \qquad (10)$$

$$Model\ C:\ y_t = \mu + \theta DU_t(\tau_b) + \beta t + \gamma DT_t(\tau_b) + \alpha y_{t-1} + \sum_{i=1}^{k} \varphi_i \Delta y_{t-i} + e_t \qquad (11)$$

where $DU_t(\tau_b) = 1$ if $t > \tau_b$ and 0 otherwise, and $DT_t(\tau_b) = t - \tau_b$ for $t > \tau_b$ and 0 otherwise. Δ is the first difference operator and e_t is a white noise disturbance term with variance σ^2. DU_t is a sustained dummy variable that captures a shift in the intercept, and DT_t represents a shift in the trend occurring at time τ_b.

Model A allows for a one - time shift in intercept; model B is a unit root test of a series around a broken trend; and model C accommodates the possibility of a change in the intercept as well as a broken trend. In applying the Zivot and Andrews (1992) test, some region must be chosen such that the end points of the sample are not included, for in the presence of the end points the asymptotic distribution of the statistics diverges to infinity (see Andrews, 1993 for details). The breakpoint is estimated by the ordinary least squares for $t = 2, 3, ..., T-1$, and the breakpoint τ_b is selected by the minimum t-statistic ($t_{\hat{\alpha}}$) on the coefficient of the autoregressive variable. $t_{\hat{\alpha}}$ is the one-sided t-statistic for testing $\alpha = 1$ in models A, B and C. We determined the lag length k using the general to specific approach adopted by Perron (1989). Given that our sample sizes are relatively small (between 32 and 46), we set $k_{\max} = 5$ and choose the order of lags such that the first t-statistic was greater than 1.6 in absolute value. The lag length is determined for each $T-2$ regressions respectively. While asymptotic critical values are available for this test, Zivot and Andrews (1992) warn that with small sample sizes the distribution of the test statistic can deviate substantially from this asymptotic distribution. To circumvent this distortion, we compute 'exact' critical values for the test following the methodology recommended in Zivot and Andrews (1992). Critical values are computed using stochastic simulations for different sample sizes T = 32, 34, 36, 39, 40, 44, 46, and 20,000 replications for the three models A, B and C. A GAUSS code is available upon request. . We reject the null of a unit root if $t_{\hat{\alpha}} < \kappa_{\inf,\alpha}$, where $\kappa_{\inf,\alpha}$ denotes the size α left-tail critical value.

EMPIRICAL RESULTS

This paper uses annual time series data on the ECOWAS countries composed of Benin, Burkina Faso, Cape Verde, Cote d'Ivoire, Gambia, Ghana, Guinea, Guinea-Bissau, Liberia, Mali, Niger, Nigeria, Senegal, Sierra Leone, and Togo. The literature suggests a considerable range of choice for measures of financial development. Sims (1972), King and Levine (1993), for example, have used monetary aggregates, such as M2 or M3 expressed as a percentage of GDP. Recently, Demetriades and Hussein (1996) and Levine and Zervos (1998) have raised doubts about the validity of the use of such variable to analyse the relationship between financial development and economic growth because GDP is a component of both focus variables (Shan and Jianhong, 2006). Moreover, Abu-Bader and Abu-Qarn (2008) underline that in developing countries, a large part of M2 stock consists of currency held outside banks. As such, an increase in theM2/GDP ratio may reflect an extensive use of currency rather than an increase in bank deposits, and for this reason this measure is less indicative of the degree of financial intermediation by banking institutions.

In this study, we use the ratio of credit to private sector to gross domestic product (GDP). The credit to private sector ratio is an appropriate measure of financial development because it is associated with mobilizing savings to facilitating transactions, providing credit to producers and consumers, reducing transaction costs and fulfilling the medium of exchange function of money (Shan and Jianhong, 2006). This indicator is frequently used in recent studies to assess the allocation of financial assets (Aghion et al. 2009, Ahlin and Pang, 2008, Bolbol et al. 2005, Baltagi et al., 2009). The series comprise yearly observations between 1960 and 2005, namely real gross domestic product per capita (denoted by GDPC) as a measure for economic growth, credit to private sector

Table 1. Zivot and Andrews (1992) unit root test results.

Countries	Variables	Time of break	Lags	$\hat{\mu}$	$\hat{\theta}$	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\alpha}$	Model type
Benin	Y	1978	1	152.18(4.49)	-21.24(-3.27)	0.41(1.85)	1.90(3.65)	-0.55(-4.45)[-6.56]	C
	F	1988	5	2.35(2.15)	-7.01(-2.84)	0.26(2.83)		-0.48(-4.33)[-6.02]	A
Burkina Faso	Y	1989	0	108.36(5.01)	-10.54(-2.96)	1.58(4.67)	0.62(-2.09)	-0.71(-4.93)[-6.07]	C
	F	1990	2	0.84(1.70)	-3.82(-4.13)	0.20(4.46)		-0.32(-4.70)[-5.74]	A
Cape Verde	Y	1990	5	200.93(2.88)		2.23(2.46)	7.21(2.78)	-0.33(-2.97)[-6.28]	B
	F	1987	3	24.40(5.90)	-4.97(-4.01)	0.07(0.79)	1.88(5.81)	-1.48(-6.08)[-6.94]	C
Cote d'Ivoire	Y	1979	0	226.63(5.51)	-152.75(-7.05)	10.75(4.39)	-13.43(-3.87)	-0.37(-3.53)[-6.59]	C
	F	1986	5	14.19(3.97)		0.77(2.76)	-1.88(-3.14)	-0.71(-3.56)[-6.28]	B
Gambia	Y	1974	0	152.08(4.27)	28.12(3.51)	-0.17(-1.00)		-0.55(-4.20)[-6.20]	A
	F	1985	0	6.95(4.17)	-8.63(-5.31)	0.32(3.25)		-0.53(-5.09)[-6.17]	A
Ghana	Y	1979	5	93.05(3.33)	-31.19(-3.42)	-0.62(-1.02)	2.43(3.33)	-0.34(-3.31)[-6.58]	C
	F	1986	0	3.57(3.20)		-0.12(-2.62)	0.39(3.43)	-0.35(-3.30)[-6.23]	B
Guinea	Y	1995	0	75.47(3.00)	10.76(3.33)	0.12(1.14)		-0.23(-3.00)[-6.32]	A
	F	1997	0	3.07(3.93)	-0.53(-2.40)	0.03(2.70)		-0.84(-4.10)[-6.40]	A
Guinea-Bissau	Y	2000	0	106.71(3.64)	-22.27(-2.38)	-0.04(-0.15)		-0.61(-3.67)[-6.27]	A
	F	1991	2	43.89(7.49)	-6.56(-5.64)	-0.02(-0.42)	-2.29(-6.71)	-2.44*(-7.48)[-6.70]	C
Liberia	Y	1980	1	184.73(4.53)	-136.88(-4.83)	-2.30(-2.81)	5.90(2.89)	-0.21(-4.35)[-6.56]	C
	F	1997	1	-5.64(-0.53)	-45.21(-2.23)	2.03(2.00)		-0.52(-3.38)[-6.48]	A
Mali	Y	1984	0	70.04(3.06)	-20.03(-2.40)	0.28(0.54)	1.33(2.28)	-0.36(-2.87)[-6.78]	C
	F	1986	4	10.00(3.36)	-4.39(-1.77)	0.08(0.81)		-0.53(-3.36)[-6.20]	A
Niger	Y	1991	4	249.84(4.23)		-4.45(-4.05)	3.82(3.23)	-0.75(-4.39)[-6.19]	B
	F	1977	4	1.39(1.48)	2.88(2.16)	0.19(1.39)	-0.38(-2.36)	-0.34(-3.78)[-6.59]	C
Nigeria	Y	1980	5	175.39(4.58)	-79.34(-4.41)	8.00(3.93)	-6.61(-3.13)	-0.61(-4.58)[-6.59]	C
	F	1992	5	2.74(2.92)	-4.09(-2.50)	0.27(3.11)		-0.65(-3.82)[-6.09]	A
Senegal	Y	1998	0	256.81(3.65)	38.50(3.34)	-1.13(-2.75)		-0.55(-3.70)[-6.04]	A
	F	1977	4	6.63(-2.98)	7.88(2.85)	0.55(2.56)	-1.00(-3.66)	-0.54(-4.75)[-6.56]	C
Sierra Leone	Y	1991	3	95.16(4.17)	-36.59(-3.38)	0.27(0.92)		-0.37(-4.23)[-6.09]	A
	F	1983	0	3.12(5.00)	-2.33(-4.17)	0.03(1.64)		-0.55(-5.52)[-6.04]	A
Togo	Y	1980	0	91.22(4.21)	-29.80(-3.41)	2.00(2.37)	-2.47(-2.15)	-0.38(-3.74)[-6.59]	C
	F	1974	0	3.35(2.02)	7.35(3.61)	0.24(1.22)	-0.46(-2.19)	-0.47(-4.84)[-6.51]	C

Notes: * denotes rejection of the null hypothesis of unit root at 5%. Numbers in (.) and [.] are respectively t-statistics and 5% critical values calculated using stochastic simulation with 20,000 replications. Y and F are related to GDP per capita and financial development indicator, respectively.

as a percentage of gross domestic product (denoted by CRE) as an indicator of financial development. Time series data are from the 2007 world development indicators of the World Bank (2007). While the bounds test for cointegration is applicable irrespective of whether the variables are integrated of order one or order zero, it is important to establish that the variables are not integrated of an order higher than one. Our second 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 the Zivot and Andrews (1992) unit root test. This test is performed on a country – by - country basis. The results for the unit root tests about GDP per capita and the ratio of credit to private sector to GDP are summarized in Table 1.

Table 1 shows that for most of the series, t - statistics are greater than the 5% critical values calculated, except for the financial development indicator in Guinea - Bissau. At the 5% level, the Zivot and Andrews test provides strong evidence that the two series ($\ln(GDPC)$ and (ln(*CRE*) have a unit root for all the ECOWAS countries, except for Guinea - Bissau where financial development has a structural breakpoint in 1991. Hence, the implementation of the Toda and Yamamoto non-causality tests requires that VAR models are augmented by one extra-lag for all ECOWAS countries. Moreover, for most of the ECOWAS countries structural breaks about economic activity appear between 1978 and 1980, corresponding to the beginning of the commodity crisis of the 1980s, while breakpoints for financial sector activities mostly occur during the period of 1985 -1990 (Figure 1) that corresponds to the start period of financial liberalization within the context of structural adjustment in the ECOWAS area. Indeed, West African countries, like

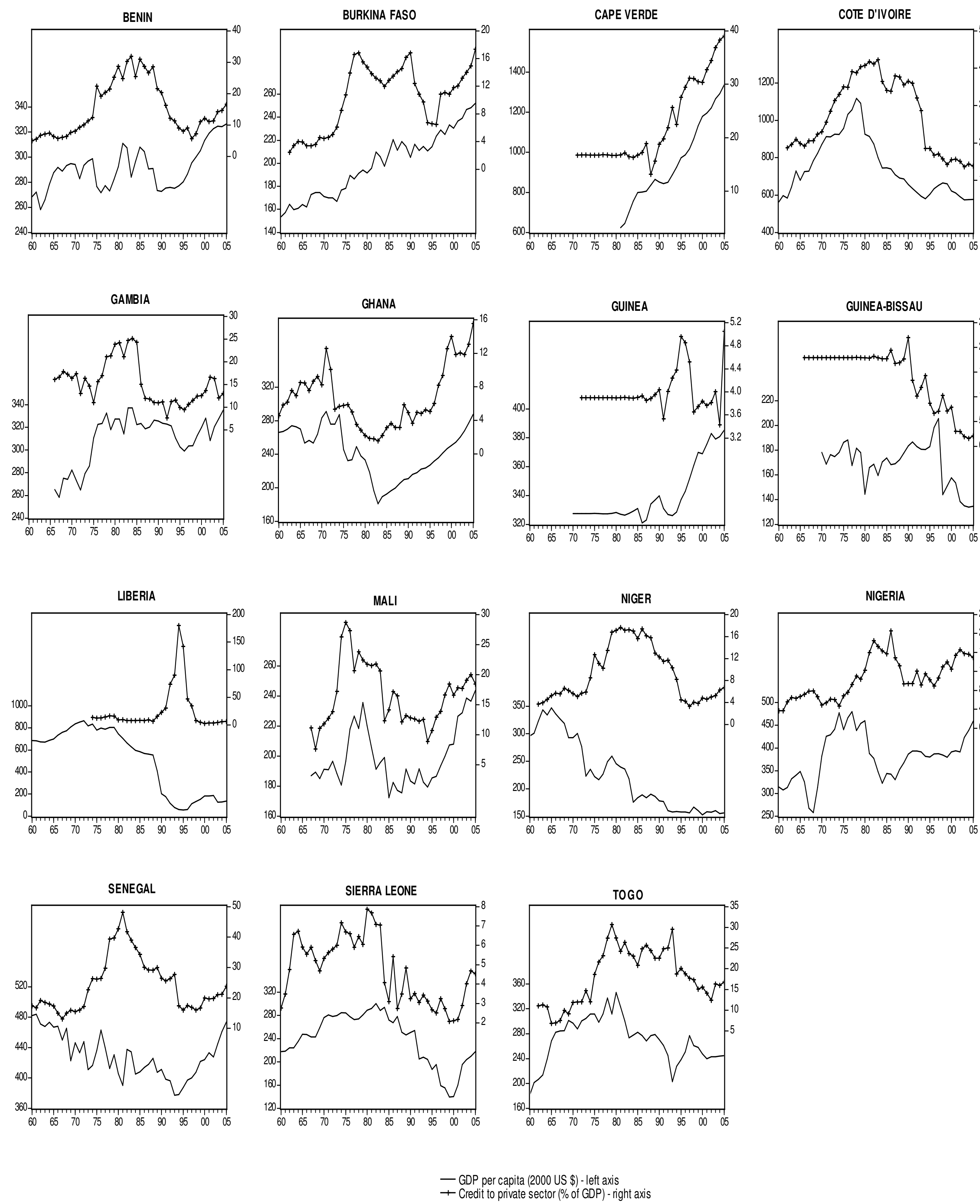

Figure 1. Annual GDP per capita and financial development dynamics of ECOWAS countries, 1960 - 2005.

Table 2. Bounds tests, F-statistics.

Countries	Samples	Dependent variable	Lags	$\chi^2(1)$	F-statistic	5% lower critical value	5% upper critical value	Cointegration
Benin	1960-2005	$Y^{\dagger}$	2	0.040	4.63	7.12	7.96	No
		$F^{\dagger}$	0	0.298	4.77	7.12	7.96	No
Burkina Faso	1962-2005	$Y^{\dagger}$	1	0.232	4.93	7.16	8.01	No
		F	1	1.958	1.83	5.27	6.17	No
Cape Verde	1981-2005	Y	1	0.006	20.50	5.44	6.35	Yes
		$F^{\dagger}$	3	1.900	2.70	7.39	8.29	No
Cote d'Ivoire	1962-2005	$Y^{\dagger}$	0	0.388	6.90	7.16	8.01	No
		F	0	0.000	8.33	5.27	6.17	Yes
Gambia	1966-2005	Y	1	1.276	3.00	5.36	6.23	No
		F	0	0.052	1.70	5.36	6.23	No
Ghana	1960-2005	$Y^{\dagger}$	1	0.111	17.29	7.12	7.96	Yes
		F	1	0.665	1.57	5.27	6.15	No
Guinea	1971-2005	Y	4	0.736	8.60	5.46	6.39	Yes
		$F^{\dagger}$	0	0.227	7.33	7.43	8.31	No
Guinea-Bissau	1970-2005	Y	0	0.174	5.79	5.42	6.33	Inconclusive
		$F^{\dagger}$	0	0.690	1.51	7.36	8.24	No
Liberia	1974-2005	$Y^{\dagger}$	3	0.738	7.07	7.50	8.42	No
		F	2	1.752	7.09	5.50	6.45	Yes
Mali	1967-2005	Y	0	0.202	6.13	5.36	6.27	Inconclusive
		$F^{\dagger}$	1	0.094	5.30	7.27	8.12	No
Niger	1962-2005	$Y^{\dagger}$	4	1.936	6.26	7.16	8.01	No
		F	0	2.419	2.49	5.27	6.17	No
Nigeria	1960-2005	$Y^{\dagger}$	1	1.585	3.69	7.12	7.96	No
		F	0	1.376	5.31	5.27	6.15	Inconclusive
Senegal	1960-2005	Y	0	0.924	2.19	5.27	6.15	No
		F	1	0.180	1.58	5.27	6.15	No
Sierra Leone	1960-2005	Y	0	1.118	4.71	5.27	6.15	No
		$F^{\dagger}$	0	0.037	7.80	7.12	7.96	Inconclusive
Togo	1962-2005	$Y^{\dagger}$	0	0.018	3.53	7.16	8.01	No
		F	0	1.430	5.28	5.27	6.17	Inconclusive

Notes: † unrestricted trend in model. $\chi^2(1)$ is an LM statistic for testing no residual serial correlation against order 1. Bounds critical values are calculated using stochastic simulation with 30,000 replications.

most other African states, entered the 1980s with a serious economic crisis which culminated in pronounced disequilibria in both the domestic and external sector. The combined effects of falling commodity prices, deteriorating terms of trade, persistent balance of payments deficits, increasing debt burdens, rapid population growth, and declining domestic output created a gloomy picture. In ought to enable economies to grow faster, economic reforms have been implemented in the ECOWAS countries, with different degrees of intensity. Financial liberalization was a significant component of these policies. Central banks liberalize interest rates, avoid or abolish the direct allocation of credit, implement monetary policy through indirect instruments and restructure and privatize banks (Reinhart and Tokatlidis, 2003). Unfortunately, many analysts of the adjustment process suggest that, in general, reforms have failed to generate real economic growth (Dorosh and Sahn, 2003) and financial reforms appear to have affected the economies in ECOWAS area very little (see also Fosu et al. 2003 for an overview of economic structural reforms in Sub-Saharan Africa).

Following the modelling approach described earlier, we determine the appropriate lag length and compute the bounds F - statistics. Akaike and Schwarz Bayesian Information criteria are used to select the optimal order of lags to include in the unrestricted error correction models. Models are estimated for m=0,1,…,5. Table 2 provides results about the bounds tests F-statistic, 5% bounds critical values and Lagrange multiplier statistics for testing the hypothesis of no residual serial correlation against order 1 denoted by $\chi^2(1)$. The Akaike and Schwarz information

Table 3. ARDL estimation.

Independent variables	Cape Verde	Cote d'Ivoire	Ghana	Guinea	Liberia
	ΔY	ΔF	ΔY	ΔY	ΔF
Long - run relationship					
Intercept	1.138* (6.940)	-1.928* (-4.027)	2.546* (5.286)	-1.176* (-3.325)	
Trend			-0.001** (-2.421)		
Y_{t-1}	-0.228* (-7.132)	0.346* (4.074)	-0.491* (-5.378)	0.146** (2.542)	0.060** (2.632)
F_{t-1}	0.142* (7.120)	-0.113** (-2.427)	0.104* (5.524)	0.241* (4.449)	-0.211* (-4.257)
Long - run effect	0.626* (9.836)	3.067* (5.713)	0.211* (7.406)	-1.646 (-9.112)	0.284* (7.173)
Short - run relationship					
EC(-1)[a]	-0.228* (-7.132)	-0.113** (-2.427)	-0.491* (-5.378)	0.146** (2.542)	-0.211* (-4.257)
ΔY_t		0.260 (1.037)			-0.883* (-2.779)
ΔY_{t-1}	0.239** (2.207)		0.323** (2.602)		-1.124* (-3.783)
ΔF_{t-1}	-0.136* (-5.233)		-0.091* (-3.462)	-0.257* (-4.325)	-0.385** (-2.138)
ΔY_{t-2}				-0.609* (-2.864)	-0.990* (-2.938)
ΔF_{t-2}				-0,114** (-2.265)	
ΔY_{t-3}				-0.296 (-1.490)	
ΔF_{t-3}				-0.139** (-2.741)	
ΔY_{t-4}				-0.488*** (-2.033)	
ΔF_{t-4}				-0.093** (-2.105)	
R - squared	0.818	0.327	0.527	0.595	0.704
$\chi^2(1)$ (p-value)	0.083 (0.773)	0.000 (0.998)	0.014 (0.905)	1.616 (0.204)	0.609 (0.435)
Observations	32	43	44	30	30

Notes: [a] EC(-1) denotes the coefficient estimate of the lagged error correction term. *, ** and *** indicate significance at the 1%, 5% and 10%, respectively. Numbers in parenthesis are t-statistics. F and Y represent natural logarithm for credit to GDP and GDP per capita, respectively. Δ is the difference operator.

criteria selected relatively small lag orders (0 or 1) for most of the countries. The $\chi^2(1)$ statistics also suggest no serial correlation against order 1 for the lag lengths selected by AIC and SC. The various F - statistics for testing the existence of a level relationship between financial development and economic growth are compared with the critical values we calculate by stochastic simulations. Models are built with or without an unrestricted trend depending on the dynamics of data. The computed F - statistic appears to be higher than the upper bound critical value at the 0.05 significance level for five models, namely, Cape Verde, Ghana and Guinea when the dependent variable is GDP per capita and, Cote d'Ivoire and Liberia when we use the credit to private sector ratio as dependent variable. Hence, for these five models the null of no long - run relationship between finance and growth is rejected. In other words, financial development and economic growth share a significant long - run relationship. The second group of countries includes those where no cointegration is found. In this panel of countries, we cannot decide whether there is a cointegrating relationship between the two variables for Guinea-Bissau and Mali when the dependent variable is GDP per capita, and for Nigeria, Sierra Leone and Togo when financial development is the dependent variable. Given these last results, causality tests may be statistically significant in these five countries. Our results support Ghirmay's (2004) conclusion about Ghana. However, cointegration results about Benin, Nigeria and Togo evidenced by Ghirmay (2004) are not confirmed in this paper. Differences between the two studies may be explained by differences in sample sizes and modeling approach. Given the findings reported in Table 2, 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 financial development on economic growth or the reverse effect, and estimates for the dynamic relationship between these two variables are provided by Table 3. The results in Table 3 indicate two groups of countries where there is a long - run equilibrium. The first one, composed of Cape Verde and

Table 4. Toda and Yamamoto non - causality test results.

Countries	Samples	F does not cause Y			Y does not cause F		
		Lags	Wald Statistics	P - value	Lags	Wald statistics	P - value
Benin	1960-2005	1	0.026	0.872	1	1.350	0.245
Burkina Faso	1962-2005	2	0.124	0.940	2	3.814	0.149
Cape Verde	1981-2005	4	65.447*	0.000	4	7.840***	0.098
Cote d'Ivoire	1962-2005	1	0.116	0.733	1	3.502***	0.061
Gambia	1966-2005	1	1.441	0.230	1	1.123	0.289
Ghana	1960-2005	2	22.392*	0.000	2	1.878	0.391
Guinea	1971-2005	3	8.885*	0.031	1	0.000	0.983
Guinea Bissau	1970-2005	1	0.017	0.896	1	0.903	0.342
Liberia	1974-2005	3	7.124***	0.068	3	5.745	0.125
Mali	1967-2005	1	5.019**	0.025	1	0.000	0.990
Niger	1962-2005	1	0.020	0.887	1	0.148	0.699
Nigeria	1960-2005	1	0.100	0.752	1	0.223	0.637
Senegal	1960-2005	1	0.000	0.999	1	2.587	0.108
Sierra Leone	1960-2005	1	5.373**	0.020	3	11.364*	0.009
Togo	1962-2005	1	0.005	0.945	1	1.168	0.280

Notes: *, ** and *** indicate significance at the 1, 5 and 10%, respectively. F and Y represent natural logarithm for credit to GDP and GDP per capita, respectively.

Ghana, is characterized by a positive and statistically significant long - run effect on GDP per capita of financial development. This long - run effect is higher in Cape Verde (0.626) than in Ghana (0.211). The second group is composed of Cote d'Ivoire and Liberia where economic growth significantly influences financial development, with a positive elasticity. The effect of GDP per capita on finance is 3.067 in Cote d'Ivoire while it is very low in Liberia (only 0.284). However, in Guinea an increase in financial development due to reforms policies is associated with low economic performance. Hence, even there is a long - run link between finance and growth, this effect is significantly negative. The long - run elasticities calculated in this study are sharply different from that shed light by Spears (1992) using data on ten African countries. Indeed, she obtains a correlation between financial development and growth close to 1. Short - run Fluctuations of financial development indicator seem to lower the GDP per capita growth rates.

The existence of a cointegrating relationship among financial development and growth for Cape Verde, Cote d' Ivoire, Ghana, Guinea and Liberia suggests that there must be causality between these variables in at least one direction. As previously mentioned, to set the stage for the Toda - Yamamoto test, the order of integration of the variables is initially determined using the Zivot - Andrews unit root test. Then, we determine the appropriate lag structures to include in the vector auto-regressive models using Akaike and Schwarz Bayesian Information Criteria. Table 4 presents the results for the non-causality from financial development to economic growth, and vice versa, in the ECOWAS countries. The fourth and seventh columns present the modified Wald statistics. We find that financial development Granger-causes economic growth in six countries: Cape Verde, Ghana, Guinea, Liberia, Mali and Sierra Leone. Hence, the result that financial development 'leads' economic growth in these six countries is consistent with the finance-led growth (or supply-leading) hypothesis previous studies by King and Levine (1993) and Levine and Zervos (1998), and can be explained by the idea that financial system liberalization enables to mobilize domestic savings. On the other hand, GDP per capita significantly causes financial development in Cape Verde, Cote d'Ivoire and Sierra Leone. These last results lend some support to the 'demand-following' view initially stated by Robinson (1952) and recently confirmed by Demetriades and Hussein (1996). In other words, economic development 'leads' to an improvement in the financial system in Cape Verde, Cote d'Ivoire and Sierra Leone. These results are also in the line of that evidenced by Spears (1992), that is causality rather runs from the GDP per capita growth rate to finance in the case of Cote d'Ivoire. However, our results are statistically stronger than Spears' because her results are improper due to a lack of stationary testing for the series. The empirical evidence provided in this study has supported the three views in the literature. We evidence (i) the 'finance - led' growth hypothesis in the case of Ghana, Guinea, Liberia and Mali, (ii) the 'demand-following' hypothesis in Cote d'Ivoire, and (iii) the bidirectional causality in the case of Cape Verde and Sierra Leone.

POLICY IMPLICATIONS

To investigate the impact on financial development and

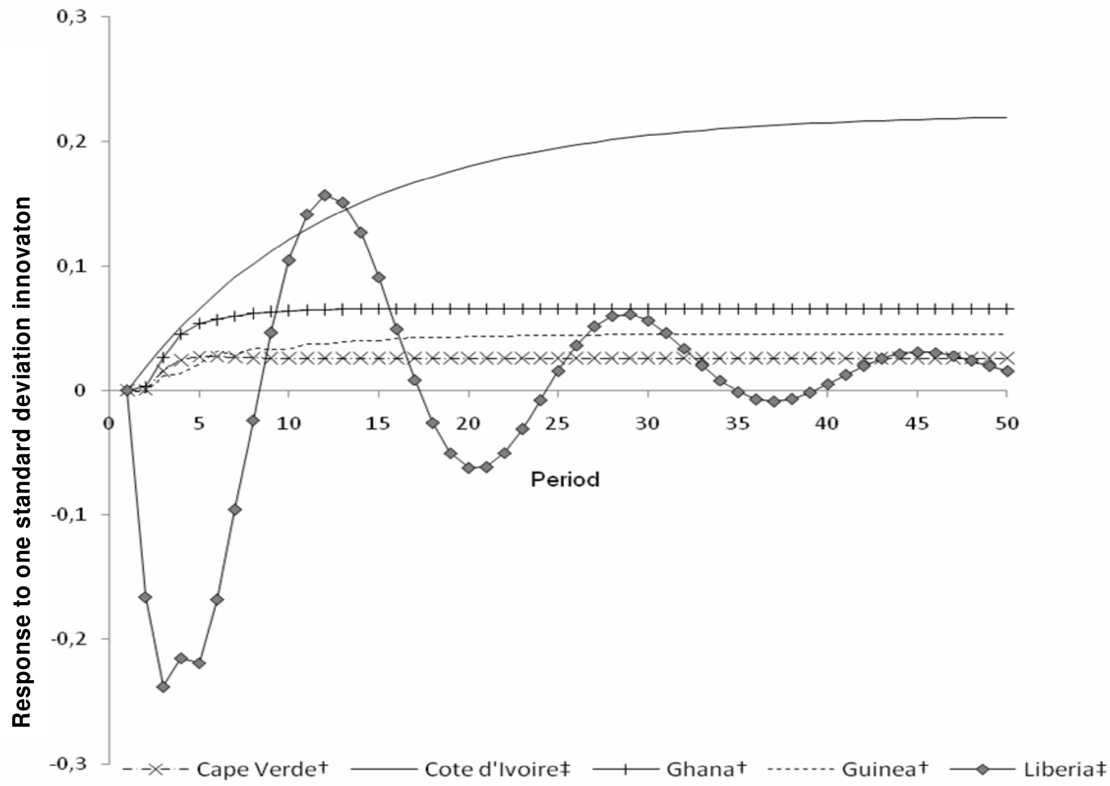

Figure 2. Response function on GDP per capita and on financial development. Notes: † denotes Response of GDP per capita to finance and ‡ Response of finance to GDP per capita.

economic growth in countries where there is a long-run relationship between these variables, we use impulse response function to trace the time paths of GDP per capita and the credit to private sector ratio in response of a one-unit shock to both finance and growth. More precisely, we simulate a positive shock to financial development in the case of Cape Verde and Ghana on the one hand and to economic growth in the case of Cote d'Ivoire and Liberia on the other hand. Figure 2 depicts the time paths of the responses of GDP per capita and financial development. It is shown that a one-unit standard deviation of financial development has a stronger and longer positive effect on economic growth in Ghana than in Cape Verde. The increase in the GDP per capita for Ghana stabilizes 11 years after the initial shock at 6.5% while in Cape Verde the increase in GDP per capita keeps up the same speed of 2.6% over 8-year horizon. This results evidence the fact that financial development, as measured by total credit to private sector divided by nominal GDP, does promote economic growth in Ghana and Sierra Leone.

Similarly, the response of a one - unit standard deviation of GDP per capita on finance confirms that economic growth also affects financial development in Cote d'Ivoire and Liberia. A shock in GDP per capita always raises the financial development indicator in Cote d'Ivoire. This effect stagnates at 2.3% after 30 years. However, the effect on finance in Liberia waves around the period - axis and converges to zero.

Finally, our results offer mixed blessing for policy makers in the ECOWAS countries. This paper provides an empirical basis for promoting financial and economic development. It has two important policy implications. First, to gain sustainable economic growth, it is desirable to further expand and improve the efficiency of the financial system through appropriate regulatory and policy reforms, and facilitate broad access to financial services, in Cape Verde, Ghana, Liberia, Mali and Sierra Leone, in order to promote faster economic growth. Second, to take advantage of the positive interaction between financial and economic development, one should promote economic growth. In other words, strategies that promote economic development in the real economy should also be emphasized, in Cape Verde, Cote d'Ivoire and Sierra Leone.

Conclusion

This study has re-examined the cointegrating and causal relationship between financial development and economic growth in the ECOWAS countries. To this end, we use two recent procedures which are the Pesaran et al. (2001) approach to cointegration and the procedure for non –

causality test popularized by Toda and Yamamoto (1995). We build unrestricted error correction models and compute bounds F - statistics to test for the absence of a long - run relationship between finance and growth. We also construct vector autoregressive models and compute modified Wald statistics to test for the non - causality from financial development to economic growth. Data are from the World Bank (2007) and cover the period 1960 - 2005. We show that there is a long - run relationship between financial development and economic growth in five countries, namely, Cape Verde, Cote d'Ivoire, Ghana, Guinea and Liberia. In addition, it is shown that GDP per capita significantly causes financial development in Cape Verde, Cote d'Ivoire and Sierra Leone. These last results lend some support to the 'demand-following' view initially stated by Robinson (1952) and recently confirmed by Demetriades and Hussein (1996). In return, financial development 'leads' economic growth in Cape Verde, Ghana, Liberia, Mali and Sierra Leone.

This conclusion is consistent with the 'finance - led' growth (or supply - leading) hypothesis previously studied by King and Levine (1993) and Levine and Zervos (1998). Our study highlights the inappropriateness of cross - sectional analysis and the necessity to examine the finance - growth nexus in a country - by - country basis because the ECOWAS countries differ in their level of financial development due to differences in policies and institutions. These results support the view of the World Bank that economic policies are country specific and their success depends on the institutions that implement them (World Bank, 1993). The findings of this paper accord with the view of other empirical studies that the relationship between financial development and economic growth cannot be generalized across countries because these results are country specific.

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