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RE-EXAMINING THE FINANCE-GROWTH NEXUS: STRUCTURAL BREAK, THRESHOLD COINTEGRATION AND CAUSALITY EVIDENCE FROM THE ECOWAS

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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. To this end, we use the Gregory and Hansen (1996a, 1996b) approach to cointegration with structural change 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 long-run relationship between financial development and economic growth in six countries, namely, Burkina Faso, Cape Verde, Cote d'Ivoire, Ghana, Liberia and Sierra Leone. In addition, we show that financial development 'leads' economic growth in Ghana and Mali while growth causes finance in Burkina Faso, Cote d'Ivoire and Sierra Leone, and a bidirectional causality in Cape Verde and Liberia. The policy implication is that Cape Verde, Ghana and Mali should give policy priority to financial reform while Burkina Faso, Cote d'Ivoire and Sierra Leone should promote economic growth.

Keywords: Threshold Cointegration, Financial Development, Granger Causality, Growth
JEL classification: C32, E44

1. 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

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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.

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), 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 (2008b), Al-Yousif (2002) or Demetriades and Hussein (1996) 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). 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 Gregory and Hansen (1996a, 1996b) approach to cointegration with structural change and the Toda and Yamamoto (1995) procedure to test for the non-causality between the variables of interest. It is now convenient in time-series analysis to check whether models chosen for describing the data under study are subject to structural breaks. When one models drifting time series, structural shifts may, in particular, influence their long-run properties. It turns out that accounting for structural breaks is also crucial for the study of integrated multivariate dynamical systems (Andrade *et al*., 2005; Kasman *et al*., 2008). The power of conventional cointegration tests falls sharply when cointegrating relationships are subject to structural changes. Lack of careful investigation of these potential structural breaks may thus lead to misspecification of the long-run properties of a dynamical system and inadequate estimation and testing procedures (Gregory *et al*., 1996).

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) 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 test for the long-run relationship instability, and test for no-cointegration with a structural change. We then build error correction models including a measure of economic growth and a financial development indicator. In addition, we construct bivariate levels vector autoregressive model and test for the non-causality from financial development to economic growth, and vice versa. Gregory and Hansen (1996a, 1996b) cointegration tests results show the existence of a long run relationship between financial development and economic growth in Burkina Faso, Cape Verde, Cote d'Ivoire, Ghana, Liberia, and Sierra Leone. Moreover, following Toda and Yamamoto (1995), there is a bi-directional causality in Cape Verde and Liberia; financial development significantly causes

[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.

economic growth in Ghana, and Mali while growth causes finance in the case of Burkina Faso, Cote d'Ivoire and Sierra Leone.

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

2. THE ECONOMETRIC FRAMEWORK

This section highlights the econometric framework used to study cointegration and causality between financial development and growth. We use the Gregory and Hansen (1996a, 1996b) cointegration approach and the Toda and Yamamoto (1995) causality testing procedure.

2.1. 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 (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: $$y_t = \mu + \theta DU_t(\tau_b) + \beta t + \alpha y_{t-1} + \sum_{i=1}^{k} \varphi_i \Delta y_{t-i} + e_t\,, \tag{1}$$

Model B: $$y_t = \mu + \gamma DT_t(\tau_b) + \beta t + \alpha y_{t-1} + \sum_{i=1}^{k} \varphi_i \Delta y_{t-i} + e_t\,, \tag{2}$$

Model C: $y_t = \mu + \theta DU_t(\tau_b) + \beta t + \gamma DT_t(\tau_b) + \alpha y_{t-1} + \sum_{i=1}^{k} \varphi_i \Delta y_{t-i} + e_t$, (3)

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 $t_{\hat{\alpha}} < k_{\inf,\alpha}$, where $k_{\inf,\alpha}$ denotes the size α left-tail critical value.

2.2. 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).

The cointegration framework of Engle and Granger (1987), and Johansen (1988) has its limitations especially when dealing with data as the Data Generating Process (DGP) may be affected by major economic events. Tests for the null of cointegration are severely oversized in the presence of structural breaks, i.e., they tend to reject the hypothesis of cointegration, albeit one with stable cointegrating parameters. In other words, the presence of structural breaks leads to inefficient estimation and therefore lower testing power, as shown by Gregory *et al*. (1996). The reason is that the residuals from cointegrating regressions capture unaccounted breaks and thus typically exhibit nonstationary behaviour. Several studies have documented the sensitivity of the outcome of the tests to structural breaks (see Wu 1998; Lau and Baharumshah, 2003; among others). One proposed approach to increase power in testing is to consider non-linear techniques instead. Several procedures have been suggested to test for cointegration in the presence of structural break. Research based on the concept of threshold cointegration includes, among others, Gregory and Hansen (1996a, 1996b), Balke and Fomby (1997), Obstfeld and Taylor (1997), Enders and Falk (1998), Enders and Granger (1998), Enders and Siklos (2001), Lo and Zivot (2001), Taylor (2001), and Hansen and Seo (2002).

In this paper, we employ a two-step error-correction model (ECM) to investigate the long-run and short run relationship between financial development and real gross domestic product. For this purpose, we make use of the Gregory and Hansen (1996a, 1996b) tests for cointegration.

The Gregory and Hansen (1996a, 1996b) tests for threshold cointegration explicitely incorporate a break in the cointegrating relationship. In fact, this approach is implemented to take into account breaks occurred in West African economies. Two steps are employed within the cointegration procedure. We first perform linearity (instability) tests in the line of Hansen (1992), as recommended by Gregory and Hansen (1996a), to determine whether the cointegrating relationship has been subject to a structural change. Hence, the three proposed tests, SupF, MeanF and L_C, are employed to verify whether the long-run relationship between finance and growth is subject to a break. The SupF test is predicated on ideas inherent in the classical Chow F-tests. The alternative hypothesis is a sudden shift in regime at an unknown point in time, and amounts to calculating the Chow F-statistic. The MeanF test is appropriate when the question under investigation is whether or not the specified model captures a stable relationship. Finally, the L_C statistic is recommended if the likelihood of parameter variation is relatively constant throughout the sample. The SupF and MeanF are calculated using the trimming region [0.15T, 0.85T], where T is sample size. As a second step, we conduct cointegration tests by allowing a break in the long-run equation, following the approach suggested by Gregory and Hansen (1996a, 1996b). The advantage of this test is the ability to treat the issue of a break (which can be determined endogenously) and cointegration altogether. The procedure offers four different models corresponding to the four different assumptions concerning the nature of the shift in the cointegrating vector: the level shift model (C), the level shift with trend model (C/T), the regime shift model (C/S) and the regime and trend shift model. To model the structural change, we define the dummy variable $D_t(T_b) = 0$ if $t \le T_b$ and 1 otherwise, where the unknown parameter T_b denotes the timing of the change point. For simplicity convenience, we present the general long-run relationship with structural break, i.e., the regime and trend shift, knowing that the three others are derived from it:

$$Y_t = \mu_1 + \mu_2 D_t(T_b) + \beta_1 t + \beta_2 t D_t(T_b) + \alpha_1 F_t + \alpha_2 F_t D_t(T_b) + \varepsilon_t \,, \tag{4}$$

where $Y = \ln(GDP)$, $F = \ln(CRE)$, and ε_t a white-noise disturbance. μ_1 and μ_2 represent the intercept before the shift and the change in the intercept at the time of the shift; β_1 and β_2 are respectively the trend slope before the shift, the change in the trend slope at the time of the shift; α_1 is the cointegrating slope coefficient before the regime shift, and α_2 denotes the change in the cointegrating slope coefficient at the time of the regime shift. The standard methods to test the null of no cointegration are residual-based. Equation (4) is estimated by ordinary least squares (OLS), and a unit root test is applied to the regression errors (Gregory and Hansen, 1996a). The time break T_b is treated as unknown and is estimated with a data dependent method, i.e., it is computed for each break point in the interval [0.15*T*, 0.85*T*], where *T* denotes the sample size (Zivot and Andrews, 1992). The date of the structural break will correspond to the minimum of the unit root test statistics computed on a trimmed sample.

If a cointegration relationship is observed between the series, Granger's Representation Theorem allows estimating the error-correction model (ECM) as follows:

$$\Delta Y_t = \gamma_0 + \gamma_1 t + \pi \varepsilon_{t-1} + \sum_{i=1}^{m} \varphi_i \Delta Y_{t-i} + \sum_{i=0}^{m} \psi_i \Delta F_{t-i} + v_t \,, \tag{5}$$

where $Y = \ln(GDP)$, $F = \ln(CRE)$, π denotes the short-run adjustment parameter, ε_{t-1} is the equilibrium error lagged one-period, and v_t a stationary process with zero mean. *m* is the lag order to include in the short-run relationship, and is selected by Akaike and Schwarz information criteria, with a maximum lag order of 6.

2.3. 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 inappropriate and can suffer from pre-testing bias. If the data are integrated but not cointegrated, then causality tests can be conducted by using the first differenced data to achieve stationarity. Granger non-causality test in an unrestricted VAR model can be simply conducted by testing whether some parameters are jointly zero, usually by a standard (Wald) F-test. Phillips and Toda (1993) show that the asymptotic distribution of the test in the unrestricted case involves nuisance parameters and nonstandard distributions. An alternative procedure to the estimation of an unrestricted VAR consists of transforming an estimated error correction model (ECM) into levels VAR form and then applying the Wald type test for linear restrictions. Toda and Yamamoto (1995) propose an interesting yet simple procedure requiring the estimation of an "augmented" VAR which guarantees the asymptotic distribution of the Wald statistic (an asymptotic χ^2 -distribution), since the testing procedure is robust to the integration and cointegration properties of the process.

We use a bivariate VAR ($p + d_{\max}$) including GDP and the credit to private sector ratio, following Yamada (1998), and examine the non-causality between these variables:

$$Y_t = \varphi_0 + \sum_{i=1}^{p} \psi_i Y_{t-i} + \sum_{i=p+1}^{p+d_{\max}} \psi_i Y_{t-i} + \sum_{i=1}^{p} \varphi_i F_{t-i} + \sum_{i=p+1}^{p+d_{\max}} \varphi_i F_{t-i} + v_{1t}, \tag{6}$$

$$F_t = \chi_0 + \sum_{i=1}^{p} \eta_i F_{t-i} + \sum_{i=p+1}^{p+d_{\max}} \eta_i F_{t-i} + \sum_{i=1}^{p} \chi_i Y_{t-i} + \sum_{i=p+1}^{p+d_{\max}} \chi_i Y_{t-i} + v_{2t}, \tag{7}$$

where $Y = \ln(GDP)$, $F = \ln(CRE)$, 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; $v_{1t} \sim N(0, \Sigma_{v1})$ and $v_{2t} \sim N(0, \Sigma_{v2})$ are the residuals of the model and Σ_{v1} and Σ_{v2} the covariance matrices of v_{1t} and v_{2t}, respectively. The null of non-causality from $\ln(CRE)$ to $\ln(GDP)$ can be expressed as:

$$H_0 : \varphi_i = 0, \qquad \forall\, i = 1,2,..., p, \tag{8}$$

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:

$$W = T(\hat{\varphi}' R' (R \hat{\Sigma}_v R')^{-1} R \hat{\varphi}), \tag{9}$$

where $\hat{\varphi}$ is the ordinary least squares estimate for the coefficient φ and $\hat{\Sigma}_v$ 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 (6) and (7). 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.

3. THE 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. 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 (2008a) underline that in developing countries, a large part of M2 stock consists of currency held outside banks. As such, an increase in the M2/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. 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 (see for example, 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 (denoted by GDP) as a measure for economic growth, credit to private sector 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).

Unit root tests are conducted 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 and the ratio of credit to private sector to

GDP are summarized in Table 1.

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	1988	2	3.34 (4.24)		0.02 (4.23)	0.01 (3.66)	-0.52(-4.22) [-6.56]	B
	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	1992	0	4.06 (4.30)	0.02 (0.66)	0.02 (4.26)	0.01 (2.77)	-0.62(-4.27) [-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	1995	5	4.41 (4.08)	0.01 (0.39)	0.05 (4.05)		-0.82(-4.05) [-6.28]	A
	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	1974	3	4.57 (4.90)	-0.06 (-1.33)	0.04 (4.23)	-0.04 (-4.08)	-0.58(-4.83) [-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, The	Y	1974	0	3.32 (5.34)		0.03 (4.63)	-0.01 (-1.29)	-0.70(-5.31) [-6.20]	B
	F	1985	0	6.95 (4.17)	-8.63 (-5.31)	0.32 (3.25)		-0.53(-5.09) [-6.17]	A
Ghana	Y	1973	5	1.27 (2.78)	0.16 (3.88)	0.01 (4.84)		-0.16(-2.72) [-6.58]	A
	F	1986	0	3.57 (3.20)		-0.12 (-2.62)	0.39 (3.43)	-0.35(-3.30) [-6.23]	B
Guinea	Y	2001	2	5.88 (3.96)	0.04 (3.02)	0.03 (4.01)		-0.78(-3.94) [-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	3.27 (4.12)	0.09 (1.04)	0.02 (3.84)		-0.70(-3.95) [-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	1988	3	2.40 (5.37)		0.01 (1.34)	0.01 (0.94)	-0.35(-5.30) [-6.56]	B
	F	1997	1	-5.64 (-0.53)	-45.21 (-2.23)	2.03 (2.00)		-0.52(-3.38) [-6.48]	A
Mali	Y	1983	3	3.55 (2.96)	0.13 (2.00)	0.02 (3.06)		-0.51(-2.95) [-6.78]	A
	F	1986	4	10.00 (3.36)	-4.39 (-1.77)	0.08 (0.81)		-0.53(-3.36) [-6.20]	A

Niger	*Y*	1990	4	6.45 (4.76)	0.08 (1.23)	0.01 (3.08)	0.02 (3.83)	-0.90(-4.76) [-6.19]	C
	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*	1979	5	3.79 (4.42)	-0.13 (-2.06)	0.01 (4.03)		-0.41(-4.52) [-6.59]	A
	F	1992	5	2.74 (2.92)	-4.09 (-2.50)	0.27 (3.11)		-0.65(-3.82) [-6.09]	A
Senegal	*Y*	1997	0	4.54 (4.11)	-0.05 (-1.34)	0.01 (4.12)		-0.61(-4.09) [-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*	1990	3	2.41 (2.96)	-0.04 (-0.46)	0.01 (1.66)	-0.01 (-1.68)	-0.38(-2.91) [-6.09]	C
	F	1983	0	3.12 (5.00)	-2.33 (-4.17)	0.03 (1.64)		-0.55(-5.52) [-6.04]	A
Togo	*Y*	1981	0	2.36 (3.51)	-0.05 (-0.78)	0.04 (2.57)	-0.03 (-2.26)	-0.41(-3.39) [-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 and financial development indicator, respectively.

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 ($\ln(GDP)$ 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 1973 and 1992, corresponding to the first oil shock and to commodity crisis of the 1980s due to the second oil shock, while breakpoints for financial sector activities mostly occur during the period of 1985-1990 (see 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 most other African states, entered the 1980s with a serious economic crisis which culminated in pronounced disequilibria in both the domestic and external sector.

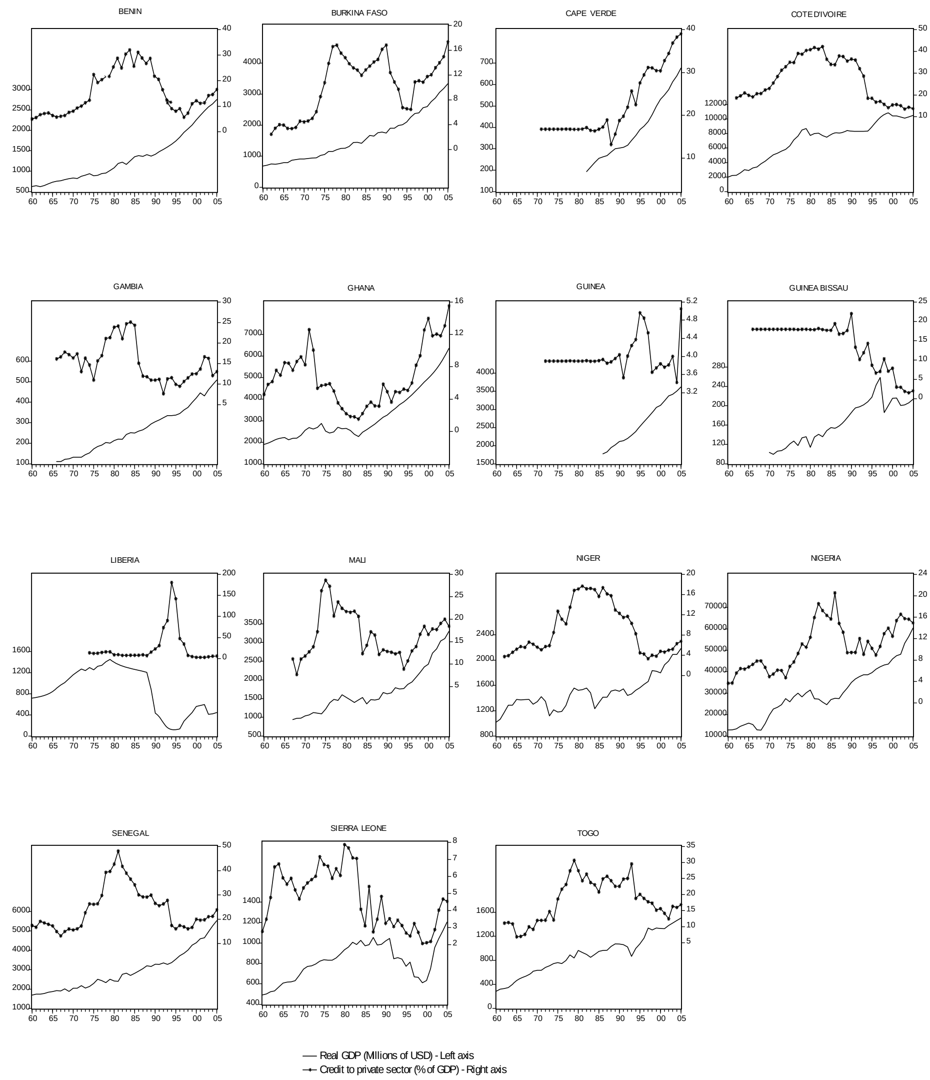

Figure 1. Annual GDP (millions of USD) and Financial Development Dynamics of ECOWAS, 1960-2005

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, 2000), 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).

Table 2. Hansen (1992) Instability Tests Results

Countries	Dependent Variable: Real GDP			Dependent Variable: Credit to Private Sector		
	SupF	*MeanF*	L_C	*SupF*	*MeanF*	L_C
Benin	1.062	0.669	0.089	2.441	1.262	0.136
	(0.200)	(0.200)	(0.200)	(0.200)	(0.200)	(0.200)
Burkina Faso	53.182*	12.990*	0.603*	0.938	0.516	0.057
	(0.010)	(0.010)	(0.042)	(0.200)	(0.200)	(0.200)
Cape Verde	152.457*	42.444*	2.559*	4.515	1.573	0.145
	(0.010)	(0.010)	(0.010)	(0.195)	(0.159)	(0.200)
Cote d’Ivoire	25.851*	6.728*	0.295	296.282*	34.096*	0.374
	(0.010)	(0.010)	(0.200)	(0.010)	(0.010)	(0.155)
Gambia, The	6.267	2.339	0.187	3.655	1.509	0.121
	(0.200)	(0.200)	(0.200)	(0.200)	(0.200)	(0.200)
Ghana	13.947**	3.910	0.147	3.502	1.470	0.166
	(0.079)	(0.200)	(0.200)	(0.200)	(0.200)	(0.200)
Guinea	1014.151*	157.272*	9.863*	9.895	3.772**	0.225
	(0.010)	(0.010)	(0.010)	(0.131)	(0.096)	(0.200)
Guinea Bissau	158143170*	6749151.6*	0.571*	3.526	0.833	0.103
	(0.010)	(0.010)	(0.049)	(0.200)	(0.200)	(0.200)
Liberia	0.584	0.413	0.067	86.106*	35.492*	0.460**
	(0.200)	(0.200)	(0.200)	(0.010)	(0.010)	(0.093)
Mali	28.888*	7.020*	0.464**	5.809	3.013	0.189
	(0.010)	(0.010)	(0.091)	(0.200)	(0.175)	(0.200)
Niger	3.320	0.959	0.123	6.967	1.759	0.142
	(0.200)	(0.200)	(0.200)	(0.200)	(0.200)	(0.200)
Nigeria	53.537*	10.330*	0.534**	10.324	2.654	0.114
	(0.010)	(0.010)	(0.060)	(0.111)	(0.200)	(0.200)
Senegal	2.666	1.304	0.172	1.410	0.655	0.057
	(0.200)	(0.200)	(0.200)	(0.200)	(0.200)	(0.200)
Sierra Leone	1.469	0.820	0.091	41.082*	23.151*	0.332
	(0.200)	(0.200)	(0.200)	(0.020)	(0.010)	(0.198)
Togo	8.663	2.686	0.320	2.436	0.613	0.072
	(0.200)	(0.200)	(0.200)	(0.200)	(0.200)	(0.200)

Notes: * and ** denote rejection of the null hypothesis of stability at 5% and 10%, respectively. Numbers in (.) are p-values.

Following the modelling approach described earlier, we first test for the instability of the long run relationship between financial development and real GDP using Hansen (1992). The test statistics *SupF*, *MeanF* and L_C are reported in Table 2.

It is shown that there is not enough evidence to reject the null of stability in five of the ECOWAS countries, namely, Benin, The Gambia, Niger, Senegal and Togo, for both long-run equations (finance and economic growth), since none of the test statistics are significant at the 10% level. Using real GDP as dependent variable, the three test statistics suggest that the long-run relationship between finance and growth may be unstable at the 10% in Burkina Faso, Cape Verde, Guinea, Guinea Bissau, Mali and Nigeria. However, when applied to the growth equation, the tests do not yield clear results in Cote d'Ivoire and Ghana. Indeed, in Cote d'Ivoire the SupF and MeanF tests lead to an unstable long-run growth equation, while in Ghana only the SupF test allows rejecting the null of stability of this equation. The finance equation seems stable in most of the ECOWAS countries, except for Liberia where the three tests lead to a conclusive unstable long-run relationship, and for Cote d'Ivoire, Guinea, and Sierra Leone where the SupF and MeanF test statistics suggest that the relationship is unstable.

As presented earlier, the next step of our modelling is the threshold cointegration tests proposed by Gregory and Hansen (1996a, 1996b). They provide an alternative approach with tests that are based on the notion of regime change and are a generalization of the usual residual-based cointegration test. These tests allow for an endogenous structural break in the cointegration. We then investigate the presence of a cointegrating relationship under structural shift between financial development and real GDP, and compute modified versions of the cointegration ADF tests of Engle and Granger (1987), as well as modified Z_t and Z_α tests of Phillips and Ouliaris (1990), i.e., $ADF^* = \inf_{T_b} ADF(T_b)$, $Z_t^* = \inf_{T_b} Z_t(T_b)$, and $Z_\alpha^* = \inf_{T_b} Z_\alpha(T_b)$. However, only ADF^* are reported in this paper for simplicity and brevity convenience. The results of the threshold cointegration tests are presented in Table 3. As reported in this table, the results of the Gregory-Hansen tests indicate rejection of the null of no cointegration in Burkina Faso, Cape Verde, Cote d'Ivoire, Liberia, and Sierra Leone, regarding the results of the Hansen (1992) instability tests. Models endogenously separate distinct regimes characterized by regime-specific parameter sets. It is shown that structural change occurs in 1970 (or 1971), 1973, 1995, and 1975 (or 1983) in the long-run growth equation for Burkina Faso, Cote d'Ivoire, Cape Verde, and Ghana, respectively. In the finance equation, structural breaks appear in 1994 (or 1988), 1996 (or 1994) and 1981 (or 1984, or 1993) for Cote d'Ivoire, Liberia and Sierra Leone, respectively. However, Table 3 suggests that cointegration exists in Guinea, Mali, Nigeria, Senegal and Togo, at the 5% level, but the error correction terms are not significant in the short-run relationship. Moreover, the Hansen (1992) instability tests do not show out the presence of structural change in the finance-growth long-run nexus in these five countries. Consequently, they will not be considered for further cointegration analysis.

Table 3. Gregory and Hansen (1996a, 1996b) Cointegration Tests Results

Countries	Dependent Variable: Real GDP				Dependent Variable: Credit to Private Sector			
	Level Shift	Level Shift with Trend	Regime Shift	Regime and Trend Shift	Level Shift	Level Shift with Trend	Regime Shift	Regime and Trend Shift
Benin	-3.62 (0) [1992]	-3.35 (0) [1998]	-3.69 (0) [1992]	-4.83 (1) [1988]	-3.42 (0) [1992]	-4.17 (0) [1989]	-3.25 (0) [1992]	-4.51 (6) [1979]
Burkina Faso	-3.18 (0) [1991]	-5.96* (0) [1970]	-3.29 (0) [1991]	-5.90* (0) [1971]	-3.10 (0) [1974]	-3.84 (0) [1971]	-3.12 (1) [1983]	-5.14 (5) [1984]
Cape Verde	-4.57 (1) [1994]	-4.54 (1) [1989]	-3.86 (0) [1999]	-5.68* (1) [1995]	-6.51* (1) [1996]	-4.34 (1) [1988]	-6.70* (1) [1996]	-5.22 (0) [1990]
Cote d'Ivoire	-4.56 (0) [1994]	-4.21 (0) [1979]	-5.15* (0) [1973]	-5.13 (0) [1980]	-4.68* (0) [1994]	-5.13* (0) [1994]	-5.41* (2) [1988]	-5.47 (0) [1992]
Gambia, The	-3.49 (0) [1986]	-4.97 (0) [1973]	-3.48 (0) [1986]	-4.90 (0) [1974]	-4.48 (0) [1986]	-4.47 (0) [1986]	-4.49 (0) [1986]	-4.44 (0) [1986]
Ghana	-4.72* (1) [1975]	-5.78* (1) [1983]	-4.56 (1) [1975]	-5.46 (1) [1983]	-4.89* (1) [1976]	-4.98 (1) [1975]	-4.80 (1) [1975]	-5.12 (1) [1974]
Guinea	-3.65 (0) [1997]	-3.55 (2) [2002]	-3.72 (0) [1997]	-3.99 (2) [2002]	-4.76* (0) [1997]	-4.74 (0) [1997]	-4.69 (0) [1997]	-5.90* (0) [1994]
Guinea Bissau	-3.23 (0) [1985]	-4.66 (0) [1998]	-3.22 (0) [1989]	-4.40 (1) [1994]	-4.85 (4) [1995]	-4.20 (0) [1978]	-4.76 (2) [1997]	-4.80 (0) [1997]
Liberia	-4.81* (0) [1996]	-5.84* (0) [1990]	-4.63 (0) [1990]	-5.72* (0) [1990]	-5.09* (0) [1996]	-6.25* (0) [1996]	-4.71 (0) [1994]	-6.83* (0) [1994]
Mali	-2.57 (0) [1989]	-4.00 (0) [1972]	-3.31 (0) [1986]	-5.20 (0) [1980]	-3.22 (0) [1984]	-4.94 (0) [1972]	-3.65 (0) [1984]	-5.55* (0) [1975]
Niger	-3.59 (4) [1995]	-4.22 (4) [1995]	-3.82 (0) [1994]	-4.67 (4) [1998]	-2.87 (4) [1972]	-3.50 (0) [1993]	-2.86 (4) [1968]	-4.20 (6) [1985]
Nigeria	-2.80 (1) [1992]	-3.82 (1) [1986]	-2.87 (6) [1980]	-5.09 (3) [1982]	-4.63* (6) [1972]	-4.22 (0) [1987]	-4.51 (6) [1972]	-3.69 (4) [1987]
Senegal	-3.07 (0) [1992]	-4.57 (0) [1998]	-2.88 (0) [1992]	-5.80* (0) [1991]	-2.75 (0) [1974]	-3.63 (4) [1988]	-3.18 (5) [1978]	-4.22 (4) [1988]
Sierra Leone	-4.17 (6) [1976]	-5.17* (3) [1992]	-4.28 (0) [1976]	-5.33 (0) [1984]	-6.07* (0) [1984]	-5.08* (0) [1993]	-5.24* (0) [1981]	-4.99 (6) [1984]
Togo	-4.49 (0) [1995]	-4.61 (2) [1968]	-5.28* (2) [1976]	-5.38 (3) [1979]	-3.84 (0) [1973]	-4.16 (0) [1973]	-4.65 (2) [1974]	-5.23 (0) [1980]

Notes: Only ADF^* is presented as threshold cointegration test statistic. * denotes rejection of the null hypothesis of no cointegration at 5%. Numbers in (.) are lag orders to include in equations. Time breaks are in [.]. 5% critical values for level shift, level shift with linear trend, regime shift, and regime and trend shift models based on Gregory and Hansen (1996a, 1996b) are respectively -4.61, -4.99, -4.95 and -5.50.

Our results support Ghirmay (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 3, we proceed with the empirical analysis only in the case of the countries where a long-run cointegrating relationship is established. We estimate short-run equations regarding according to the significant long-run relationships for a given country where threshold cointegration is evidenced. Then, we finally consider the error-correction model on the basis of Akaike and Schwarz information criteria, and results of residuals tests.

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 4. The results in table 4 indicate diverse situations in the ECOWAS countries where there is a long-run equilibrium. Indeed, Ghana is characterized by a positive and statistically significant long-run effect on GDP of financial development. In Burkina Faso, this relationship is positive and significant in the second regime occurring in 1971, while finance positively impacts real GDP before 1973. This long-run effect is higher in Cote d'Ivoire (1.614) than in Ghana (0.187) and Burkina Faso (0.048). However, in Cote d'Ivoire an increase in financial development due to reforms policies is recently (after 1973) associated with low economic performance. Hence, even there is a long-run link between finance and growth, this effect is significantly negative in Cote d'Ivoire in recent years. In Cape Verde, financial development negatively impacts real GDP in both regimes, but the effect of the second regime (after 1995) is deeper (-0.397). Moreover, Liberia and Sierra Leone are characterized by opposite figures. Indeed, the effect of real GDP on financial development is negative and significant in both regimes (the breakpoint is 1994) in Liberia, while it is positive and significant at 1% in Sierra Leone.

Table 4. ECM Estimation results

Independent variables	Burkina Faso	Cape Verde	Cote d'Ivoire	Ghana	Liberia	Sierra Leone
	ΔY	ΔY	ΔY	ΔY	ΔF	ΔF
	T_b=1971	T_b=1995	T_b=1973	T_b=1983	T_b=1994	T_b=1993
			Long-run Relationship			
μ_1	6.644*	4.566*	3.202*	7.177*	13.818*	-7.628*
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
μ_2	-0.336*	0.285	6.786*	-0.085**	2.063	0.345***
	(0.000)	(0.653)	(0.000)	(0.015)	(0.218)	(0.065)
β_1	0.032*	0.050*		0.024*	-0.056*	-0.041*
	(0.000)	(0.000)		(0.000)	(0.001)	(0.000)
β_2	0.004	0.019***			0.021	
	(0.336)	(0.064)			(0.621)	
α_1	-0.098	-0.121***	1.614*	0.187*	-1.431*	1.486*
	(0.149)	(0.068)	(0.000)	(0.000)	(0.000)	(0.000)
α_2	0.146**	-0.276	-1.898*		-0.650**	
	(0.037)	(0.331)	(0.000)		(0.025)	

F-statistics	2818,2*	669.48*	216.91*	506.78*	80.83*	25.57*
(p-value)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
			Short-run Relationship			
EC(-1)[a]	-0.964*	-0.559*	-0.180**	-0.546*	-1.345*	-0.583*
	(0.000)	(0.001)	(0.011)	(0.000)	(0.008)	(0.000)
Constant	0.017**	0.028*	0.081*	0.012***	-0.089	
	(0.046)	(0.003)	(0.000)	(0.073)	(0.165)	
Trend	0.001***		-0.002*			
	(0.074)		(0.004)			
ΔY_t					-1.613*	0.755**
					(0.000)	(0.040)
ΔF_t	0.072*	-0.091*	0.044***	0.053*		
	(0.009)	(0.003)	(0.077)	(0.042)		
ΔY_{t-1}		0.573*		0.507*		
		(0.000)		(0.001)		
ΔF_{t-1}		-0.042		-0.059**		
		(0.146)		(0.032)		
R-squared	0.451	0.719	0.382	0.411	0.520	0.327
χ^2 (1)	0.745	0.408	1.270	0.0004	1.483	0.014
(p-value)	(0.388)	(0.523)	(0.259)	(0.984)	(0.223)	(0.905)
Observations	43	22	43	44	31	45

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 p-values. *F* and *Y* represent natural logarithm for credit to GDP and GDP, respectively. Δ is the difference operator. Long run and short run equations are respectively $Y_t = \mu_1 + \mu_2 D_t(T_b) + \beta_1 t + \beta_2 t D_t(T_b) + \alpha_1 F_t + \alpha_2 F_t D_t(T_b) + \varepsilon_t$, and $\Delta Y_t = \gamma_0 + \gamma_1 t + \pi\varepsilon_{t-1} + \sum_{i=1}^{p} \varphi_i \Delta Y_{t-i} + \sum_{i=0}^{p} \psi_i \Delta F_{t-i} + v_t$, where T_b is the time break and $D_t(T_b) = 1$ if $t > T_b$ and 0 otherwise.

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 growth rates in Cape Verde, while improvement of financial sector in Burkina Faso, Cote d'Ivoire and Ghana tends to increase real GDP.

The existence of a cointegrating relationship among financial development and growth for Burkina Faso, Cape Verde, Cote d'Ivoire, Ghana, Liberia and Sierra Leone 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 autoregressive models using Akaike and Schwarz Bayesian Information Criteria. Table 5 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, Liberia and Mali. Hence, the result that financial development ‘leads’ economic growth in these four 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 significantly causes financial development in Burkina Faso, Cape Verde, Cote d’Ivoire, Liberia 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 these five ECOWAS countries. These results are also in the line of that evidenced by Spears (1992), that is, causality rather runs from the GDP 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.

Table 5. Toda and Yamamoto Non-Causality Test Results

Countries	Samples	*F* doesn’t cause *Y*			*Y* doesn’t cause *F*		
		Lags	Wald Statistics	P-value	Lags	Wald Statistics	P-value
Benin	1960-2005	1	0.361	0.550	1	0.519	0.471
Burkina Faso	1962-2005	2	0.103	0.949	2	4.857**	0.045
Cape Verde	1981-2005	4	69.098*	0.000	4	11.013**	0.026
Cote d’Ivoire	1962-2005	1	0.461	0.497	1	3.895**	0.048
Gambia	1966-2005	1	2.750	0.097	1	0.631	0.427
Ghana	1960-2005	3	15.379*	0.001	3	1.206	0.751
Guinea	1971-2005	1	0.214	0.643	1	0.077	0.780
Guinea Bissau	1970-2005	1	0.368	0.544	1	0.134	0.712
Liberia	1974-2005	3	7.841**	0.047	3	7.073***	0.069
Mali	1967-2005	1	4.172**	0.041	1	0.154	0.695
Niger	1962-2005	1	0.086	0.769	1	0.000	0.992
Nigeria	1960-2005	1	0.369	0.543	1	1.052	0.305
Senegal	1960-2005	1	0.092	0.762	1	1.538	0.215
Sierra Leone	1960-2005	3	4.092	0.252	3	9.136**	0.027
Togo	1962-2005	1	0.625	0.429	1	0.703	0.402

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

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 and Mali, (ii) the ‘demand-following’ hypothesis in Burkina Faso, Cote d’Ivoire and Sierra Leone, and (iii) the bidirectional causality in the case of Cape Verde and Liberia.

4. CONCLUSION AND POLICY IMPLICATIONS

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 Gregory and Hansen (1996a, 1996b) approach to cointegration with structural change and the procedure for non-causality test popularized by Toda and Yamamoto (1995). We test for the instability of the long-run relationship between finance and growth and test for cointegration in presence of breakpoint. 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 with structural break between financial development and economic growth in six countries, namely, Burkina Faso, Cape Verde, Cote d'Ivoire, Ghana, Liberia, and Sierra Leone. In addition, it is shown that real GDP significantly causes financial development in Burkina Faso, Cote d’Ivoire, Cape Verde, Liberia 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, and Mali. 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.

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 and Mali, 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 Burkina Faso, Cape Verde, Cote d'Ivoire and Sierra Leone.

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