Empir Econ (2013) 45:987–1008
DOI 10.1007/s00181-012-0633-x

Democracy and economic growth in Sub-Saharan Africa: a panel data approach

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Received: 16 July 2011 / Accepted: 16 July 2012 / Published online: 12 September 2012

Abstract This paper studies the link between democracy and economic development for 28 countries of Sub-Saharan Africa for the period 1980–2005 in a panel data framework. A democracy index constructed from the Freedom House indices. A variety of panel data unit root and cointegration tests are applied. The variables are found to be integrated of order one and cointegrated. The Blundell–Bond system generalized methods-of-moments is employed to conduct a panel error-correction mechanism based causality test within a vector autoregressive structure. Economic growth is found to cause democracy in the short-run, while bidirectionality is uncovered in the long-run. In addition, the long-run coefficients are estimated through the panel fully modified ordinary least squares and dynamic ordinary least squares methods. Democracy has a positive impact on GDP and vice versa. These results lend support to the virtuous cycle hypothesis.

Keywords Democracy · GDP · SSA · Panel data

JEL Classification C33 · O40

1 Introduction

Despite its huge resource endowments, the Sub-Saharan Africa (SSA), which adds up to 48 of the 54 African countries, remains to date, the poorest region in the world. In 2007, its gross domestic product (GDP) per capita was estimated to be only about \$1,869 (World Bank 2009). The SSA has been severely marked by a long series of civil wars and political turmoil (Jézéquel 2006). Democracy has been lacking and this

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Table 1 GDP and democracy trend, 1980–2005

Countries	Real GDP (million $)					Democracy index				
	1980	1990	2000	2005	1980–2005	1980	1990	2000	2005	1980–2005
Benin	1084.3	1411.8	2254.8	2727.2	1738.7	1.5	3.0	6.0	6.0	3.9
Burkina Faso	1100.6	1555.6	2610.9	3529.5	1972.3	2.5	2.5	4.0	4.0	3.0
Burundi	559.4	865.1	709.1	789.7	749.4	1.5	1.5	2.0	4.0	1.8
Cameroon	6338.8	8792.7	7990.3	12056.5	9277.9	2.0	2.0	2.0	2.0	2.0
Cen. Af. Rep.	735.3	814.6	959.4	914.8	836.4	2.0	2.5	4.5	3.5	2.9
Chad	664.9	1105.7	1385.1	2776.4	1279.3	1.5	1.5	2.5	2.5	2.0
Congo, DR.	7015.8	7659.5	4305.8	5238.6	6185.9	1.5	2.0	3.0	3.0	2.6
Congo, Rep.	1746.4	2795.6	3219.9	3975.5	2919.9	2.0	2.0	1.5	2.0	1.7
Cote d'Ivoire	7727.4	8297.6	10417.0	10409.0	8961.3	2.5	3.0	2.5	2.0	2.6
Gabon	3594.3	4298.5	5067.8	5523.0	4578.8	2.0	4.0	3.5	3.0	3.0
Gambia, The	213.5	304.5	420.9	508.5	333.7	5.5	6.0	2.0	3.5	4.0
Ghana	2639.9	3266.9	4977.5	6364.1	3852.1	5.5	2.5	5.5	6.5	3.7
Kenya	7086.8	10557.3	12705.3	15160.4	10721.5	3.5	2.0	2.5	5.0	2.7
Liberia	1390.9	433.0	560.9	444.2	699.3	2.0	1.0	2.5	4.0	2.5
Madagascar	3098.7	3265.6	3877.6	4339.1	3308.9	2.0	4.0	5.0	5.0	4.0
Malawi	999.8	1243.0	1743.5	1827.1	1359.9	1.5	1.5	5.0	4.0	3.1
Mali	1536.0	1630.1	2422.5	3294.1	1973.2	1.5	2.5	5.5	6.0	3.8
Mauritius	1518.4	2679.0	4469.3	5474.6	3204.0	5.0	6.0	6.5	7.0	6.3
Niger	1523.4	1507.0	1798.4	2207.0	1636.9	1.5	2.5	4.0	5.0	2.8
Nigeria	31451.8	34977.8	45983.6	60557.0	38045.5	5.5	3.0	4.0	4.0	3.2
Rwanda	1457.0	1781.6	1810.9	2351.1	1701.1	2.0	2.0	1.5	2.5	1.9
Senegal	2682.7	3463.3	4691.8	5891.1	3871.6	4.0	4.5	4.5	5.5	4.4
Sierra Leone	935.3	1021.6	633.8	1201.6	913.6	3.0	2.5	3.5	4.5	3.0
South Africa	95502.5	110944.7	132877.6	160792.9	117675.6	2.5	3.5	6.5	6.5	4.5
Swaziland	554.0	1023.8	1388.7	1561.6	1052.5	3.0	2.5	2.5	2.0	2.5
Togo	964.2	1070.7	1329.1	1479.5	1121.9	1.5	2.0	3.0	2.5	2.3
Zambia	2729.8	3027.5	3237.7	4088.7	3101.0	2.5	2.5	3.5	4.0	3.5
Zimbabwe	4376.4	6733.7	7399.2	5618.1	6382.0	4.5	3.0	2.5	1.5	2.9

Source Computed. *Note* In connection with the Democracy index, countries whose ratings average between 1.0 and 2.9 are Not Free, those between 3.0 and 5.4 are Partly Free, and those between 5.5 and 7.0 are Free

has also arguably led to low income levels in the region. As stated by Diamond (2005), "… *Africa cannot develop without democracy and that democracy in Africa ultimately cannot be sustained without development*." Furthermore, he defines democracy "… *as a system of government in which the people choose their leaders and representatives, and can replace them, in regular, free and fair elections*." As exposed in Table 1, while several states of the SSA are still undemocratic (Not Free), for some, there has been a slow but firm march towards democracy. From an economic perspective, the question of whether democracy affects economic growth is critical to policymakers who seek to promote greater political freedom to fuel economic development in the SSA.

This paper explores the linkage between democracy and economic growth for a sample of 28 countries of the SSA[1] over the period 1980–2005. Two generations of panel unit root and cointegration tests are conducted. These are followed by a panel vector error-correction mechanism (VECM)-based causality test, utilizing the system generalized methods-of-moments (GMM) technique. The long-run estimates are computed by means of the panel fully modified ordinary least squares (FMOLS) and dynamic ordinary least squares (DOLS) estimators. Democracy scores are obtained from the Freedom House website. Real GDP (constant 2000) data are obtained from the World Bank CD-ROM 2008. Similar to Narayan et al. (2011), real GDP is used to capture economic growth. The democracy index is simply built by averaging of the sum of civil liberties and political rights indices.[2] The political rights index shows how fair and meaningful elections are carried out, while the civil liberties index involves freedom of press, freedom of speech, freedom of religious belief, and the right to protest and organize. These scores are re-adjusted whereby a score of 1 means least free, whereas a score of 7 reflects most free.

Freedom House data are quite popular and have been used extensively in the literature. Such application has to do with a variety of studies examining the implications of economic freedom on economic growth (Hanke and Walters 1997), political institutions on the environment (Bhattarai and Hammig 2001), financial development on corruption (Altunbaş and Thornton 2012), among others. In addition to the Freedom House, other freedom indicators are compiled by the Fraser Institute, Heritage Foundation, International Institute for Management Development (IMD), and World Economic Forum. These are assessed and compared by Hanke and Walters (1997), who pinpoint to several limitations accruing during the construction of the indicators by those institutions. However, as argued by Hanson (2003), "… *Freedom House's renown as an arbiter of political freedom and civil liberties gives its index instant credibility, which may explain the failure of Hanke and Walters to challenge it* (p. 642)."

The Freedom House index does have its criticisms. According to Minier (1998), the subjectivity involved when building the index brings in some measure of error and bias. Democracy is a multifaceted theme and the index is argued to be based on a checklist which includes limits on suffrage, freedom of the press, and restrictions on individuals running for office. In addition, the overall ranking done by the Freedom House can be debated to be entirely impressionistic. The index also tends to force a seemingly continuous variable into a discrete ranking system. Nevertheless, as argued by Tavares and Wacziarg (2001), the index is fashioned with the intension of consistency across time and across countries and this makes it suitable to use in a panel data setting. The rest of the paper is set out as follows: Section 2 provides an overview of the theoretical and empirical literature involving the democracy–economic growth nexus. Section 3 presents the testing frameworks. Section 4 discusses the results. Section 5 concludes.

[1] The selection of countries is done purely on the availability of data.

[2] The characteristic of each score is available online at http://www.democracyweb.org/about/fiw1.php.

2 Review of theoretical and empirical literature

Since the seminal work of Lipset (1959), a voluminous number of scholarly works analyzing the link between economic development and democracy has emerged. As stated by him, "*perhaps the most widespread generalization linking political system to other aspects of society has been that democracy is related to the state of economic development*," where educated people are apt to "…*believe in democratic values and support democratic practices* (p. 75)." On the word of Pennar et al. (1993), "… *rising incomes at first go toward needed goods and investment, then later toward more and more of what economists call "luxury goods," such as higher education. A more educated population tends to demand political and civil rights, and so democratization begins*." Economic growth is therefore conducive to democracy (Huber et al. 1993; Barro 2002). The Lipset hypothesis will be supported if in the long-run, a change in economic growth causes a change in democracy and a rise in economic growth results in greater democracy.

The impact of democracy on economic growth is less straightforward and has been a matter of much more controversy among scholars. Mixed findings have been observed.[3] To illustrate this connection, Sirowy and Inkeles (1990) put forward three major perspectives. These are namely the "*conflict*," "*compatibility*," and "*skeptical*" ones. First, the conflict perspective views democracy as a big hurdle to economic growth. To experience economic expansion, policies inhibiting excessive increase in real wages and promotion of both national and foreign capital accumulation are required. Unless these policies are adopted, rapid growth of industrialization has a tendency to to be delayed. This in turn slows down the process of economic growth. Democratic governments are looked upon to be vote maximizers and they are more concerned about implementing myopic policies such as state benefits and welfare policies at the expense of accumulation. Democracy has thus a negative impact on economic growth (Tavares and Wacziarg 2001; Heo et al. 2008). In essence, the conflict hypothesis will be supported if in the long-run, a change in democracy causes economic growth and a rise in democracy has a negative effect on economic growth.

On the contrary, the compatibility perspective is incompatible with an authoritarian model where economic development is exclusively directed by a centralized body or dictator. As a consequence, democracy is considered to sustain equitable allocation of resources and power, reduce distributional conflicts, and support fundamental civil liberties and political rights. These are appropriate to create the necessary socio-politico-economic conditions favorable to economic growth. The impact of democracy on economic growth is expected to be positive (Kurzman et al. 2002; Ghosh and Gregoriou 2009). Fundamentally, the compatibility hypothesis will be supported if in the long-run, a change in democracy causes economic growth and an increase in democracy has a positive effect on economic growth.

[3] Kurzman et al. (2002) review 47 quantitative studies of the effect of democracy on economic growth. 19 found a positive relationship between democracy and growth, 6 found a negative relationship, and 10 reported no statistically significant relationship. 7 studies found a combination of positive and non-significant results, depending on the model used and the cases included; 2 found a combination of negative and non-significant results; 2 found mixed positive and negative results; and 1 (Barro 1996) reported an inverted-U effect.

Finally, from the skeptical perspective, there is no systematic relationship between democracy and economic growth (Rodrik 1997). With regard to this perspective, factors such as the effectiveness of government policies, institutional maturity, and the coordination of government entities, etc. play a more significant role in economic performance than the presence or absence of democracy (Heo 2010). For instance, Przeworski and Limongi (1993) find no conclusive effect of democracy on economic growth. As stated by them "… *the impact of political regimes on growth is wide open for reflection and research* (p. 66)." The skeptical perspective will be supported if no causal link between democracy and economic growth exists. In addition, a symbiotic link between democracy and economic growth can correspondingly exist (Bhalla 1997), with both having a positive effect on each other. This can be referred to the virtuous cycle hypothesis. The Lipset and compatibility hypotheses are assumed hold at the same time.

The existing empirical literature[4] has largely revolved around correlation and regression techniques while the causal effect between democracy and economic growth has actually been overlooked. As discussed above, reverse causality is an obvious possibility and this can lead to biased estimates. So far, only a few studies have tried to examine any causal link between those two variables and these are mainly based on time series analysis. Heo and Tan (2001) study such link for 32 developing countries over the period 1950–1982. Their results support a causal relationship running from economic growth to democracy for Costa Rica, Egypt, Guatemala, India, Israel, Mexico, Nicaragua, South Korea, Thailand, Uruguay, and Venezuela. Next, unidirectional causality running from democracy to growth is found for Bolivia, Burma, Colombia, Ecuador, El Salvador, Indonesia, Iran, Paraguay, Philippines, and South Africa. Bidirectional causality is stumbled upon for Chile, Dominican Republic, and Turkey. They find no relationship between economic growth and democracy for Argentina, Brazil, Haiti, Honduras, Pakistan, Panama, Peru, and Sri Lanka.

Using ECM techniques, a few studies have applied more sophisticated causality tests to examine any short-run or long-run effects. Narayan and Smyth (2006) find a bidirectional causal relationship between democracy and economic growth in the short-run. A unidirectional causal relationship running from democracy to economic growth is likewise discovered in the long-run for China over the period 1972–1999. Narayan et al. (2011) examine similar link for 30 countries of SSA over the period 1972–2001. They use two democracy indices, viz. the Freedom House political rights index and the Legislative index of Electoral Competitiveness (LIEC). They find support for the Lipset hypothesis for Botswana and Niger with both datasets, for Chad with the Freedom House data and for Cote d'Ivoire and Gabon with the LIEC data. Support for the compatibility hypothesis is found for Botswana with the Freedom House data and for Madagascar, Rwanda, South Africa, and Swaziland with the LIEC data. Support for the conflict hypothesis is found for Gabon with the Freedom House

[4] A relevant study includes Burkhart and Lewis-Beck (1994) where they examine the causal link between economic development and economic growth for 131 countries over the period 1972–1989. They locate unidirectional causality running from economic development to democracy. Campos and Nugent (2002) test whether a causal and negative long term relation exist between political instability and economic growth 98 countries and find no evidence of such relationship.

data and Sierra Leone with the LIEC data. For most of the countries in their sample, the skeptical hypothesis is supported.

One major problem with econometrics research is the lack of sufficient observations over long periods in the context of the need to test for unit roots and cointegration amongst the series. Failure to do so exposes the study to the criticism of spurious results. Time series tests can be subject to criticism of low power especially when the number of observations is relatively small. As claimed by Toda (1995) "… *100 observations are not sufficient to ensure good performance of the tests* (p. 79)." One potential solution to this problem is to employ panel data techniques. These allow for a significant increase in testing power relative to conventional time series methods (Coakley and Fuertes 1997). This paper attempts to add new empirical evidence to the literature by revisiting the democracy–economic nexus for the SSA within a panel data framework.

3 Testing framework

To test for any causal effects between democracy and economic growth, a panel Granger-type causality test is constructed. Granger causality does not however mean true causality. Such statistical concept is based on linear regression stochastic processes (Granger 1969). It can only be interpreted as showing whether prior changes in one series add (or do not add) significantly to the explanation of the future value of another series (Farr et al. 1998). The Granger causality test requires variables to be stationary, as it can yield spurious results (Granger and Newbold 1974). This problem can be avoided by taking advantage of a VECM-based causality test. Some preliminary testing such as unit root and cointegration tests are required before carrying out the Engle–Granger residual-type causality approach. First, both variables have to be non-stationary and integrated of same order, such as order one, denoted as I(1), to test for cointegration (Engle and Granger 1987). Macroeconomic variables tend to be non-stationary in nature. A series Y_t is integrated order of d, i.e., $Y_t \sim \mathrm{I}(d)$, if it were to be differenced by d times to become stationary. Cointegration between democracy and real GDP should be established before proceeding to the VECM-based causality test.

To verify the order of integration, a battery of panel unit root tests will be performed. Inferences on a single test can lead to inappropriate conclusion about the order of integration as none of the current panel unit root tests are devoid of statistical limitations in terms of size and power properties. The first generation tests are conceptualized by Hadri (2000), Levin et al. (2002, LLC), Im et al. (2003, IPS), and Im et al. (2005, ILT). These tests do not explicitly control for cross-sectional dependence. Each panel is assumed to operate independently from each other. This condition is rather unrealistic. Recent second generation relaxes the assumption of cross-sectional independence and account for any correlation among the panels. These tests are derived by Pesaran (2007); Chang and Song (2009) and Hadri and Kurozumi (2012).

Unit root tests are usually worked out using two distinct specifications. One test includes a constant term only and the other contains both a constant term and a time trend. These tests can be sensitive to the inclusion or exclusion of a trend. There should be a careful use of a deterministic trend or "*otherwise results can be misleading* (p. 51)" (Ahking 2002). Macroeconomic data tend to exhibit a trend over time. Intuitively, it is more apt to consider a regression with a constant and a trend at level form. Since first-differencing tends to remove any deterministic trends in the variables, inferences will be done in line with a specification with a constant term only. For the sake of completeness, both specifications are computed.

After checking the order of integration of the two variables, panel cointegration tests can be afterward conducted. First, Nyblom and Harvey (2000) propose a non-parametric cointegration test of common trends. The test assumes where H_0 is the stationarity of the series around a deterministic trend and there exists $k < n$ common trends (where rank $(\Sigma\,\eta) = k$), against the alternative of a random-walk component incidence, where there exists more than k common trends (where rank $(\Sigma\,\eta) > k$). This test tests for the H_0 of 0 common trends against the hypothesis of common trends among the variables. The non-parametric cointegration test is followed by two parametric ones. Second, Pedroni (1999) proposes several tests with the H_0 of no cointegration. These computed statistics are called within-dimension- and between-dimension-based statistics. Similar to panel unit root tests, panel cointegration tests can suffer from the problems of cross-sectional dependence. Westerlund (2007) advocates a cointegration test which can handle this issue. This can be considered as a second-generation panel cointegration test. This test is based on structural instead of residual dynamics and there is no one common factor restriction. Baltagi (2008) provides a thorough review of the panel unit root and cointegration literature which indeed points towards the vital importance of controlling for cross-sectional dependence.

If the variables are cointegrated, causality should occur in at least one direction (Baffes and Shah 1994). Consistent with Jaunky (2011), the ρth order VECM structure can be represented as follows:

$$\begin{bmatrix} \Delta\text{LGDP}_{it} \\ \Delta\text{LDEM}_{it} \end{bmatrix} = \begin{bmatrix} \alpha_1 \\ \alpha_2 \end{bmatrix} + \sum_{k=1}^{\rho} \begin{bmatrix} \beta_{11k} & \beta_{12k} \\ \beta_{21k} & \beta_{22k} \end{bmatrix} \begin{bmatrix} \Delta\text{LGDP}_{it-k} \\ \Delta\text{LDEM}_{it-k} \end{bmatrix} + \begin{bmatrix} \phi_1 \\ \phi_2 \end{bmatrix} \left[\text{ECM}_{it-1}\right] + \begin{bmatrix} \varepsilon_{1it} \\ \varepsilon_{2it} \end{bmatrix} \quad (1)$$

where $i = 1, \ldots, N$; $t = \rho + 1, \rho + 2, \ldots, T$; the αs, β_ks, and ϕs are parameters to be estimated. ECM_{it-1} represents the one period lagged error-term derived from the cointegrating vector and the error terms ε_{1it} and ε_{2it} are assumed to be serially independent with mean zero and finite covariance matrix. Given the use of a vector autoregressive (VAR) model, all variables are treated as being endogenous. LDEM_{it} and LGDP_{it} denote the natural logarithmic of real GDP and democracy for country i ay time t. A simple Wald test for the joint significance can be applied to examine the direction of any causal relationship. For instance, democracy does not Granger-cause economic growth if and only if all the coefficients β_{12k}; $\forall = 1, 2, \ldots, \rho$ are not significantly different from zero in Eq. (1). The dependent variable reacts merely to short-term shocks. In the similar fashion, economic growth does not Granger-cause democracy in the short-run if and if all the coefficients β_{21k}; $\forall = 1, 2, \ldots, \rho$ are not significant from zero. These can be referred to as the "*short-run Granger causality*" test. The coefficients on the *ECM*s represent how fast deviations from the long-run equilibrium are eliminated. An additional channel of causality can be studied by testing the significance of the *ECM*s. This test is referred to as the "*long-run Granger causality*" test.

The estimators in Eq. (1) are biased due to the correlation among the lagged dependent variables and the error terms. To account for the correlation and endogeneity problems, Arellano and Bond (1991) suggest a two-step difference GMM approach, where the lags of explanatory variables in levels are to be used as instruments. In the first step, the error terms are assumed to be independent and homoscedastic. In the second step, the first-step residuals are applied to construct consistent variances and covariances matrices, with the former assumptions are relaxed. For the instruments to be valid, ε_1 and ε_2 should not be serially uncorrelated. This condition occurs when there is statistical evidence of a significant and negative first-order serial correlation (AR(1)) and no evidence of second-order serial correlation (AR(2)) in the differenced residuals mutually. The optimal lag length ρ is chosen if such condition is satisfied.

The Arellano–Bond estimator suffers from a lack of power of the internal instruments. Following Blundell and Bond (1998), if the lag of the dependent variable and the explanatory variables are persistent in time, lags of the levels of these variables will be weak instruments for the equation in differences. They consequently recommend the system GMM. This estimator is a linear combination of the levels and differences, where the weight specified to the levels estimators grows in the prevalence of weak instruments owing to the high persistency in the series. In the presence of heteroskedasticity and serial correlation, the two-step system GMM employs of a consistent estimate of the weighting matrix, using the residuals from the one-step estimator (Davidson and MacKinnon 2004). In contrast to the conventional OLS method, the system GMM does not assume normality and controls for heteroskedasticity.

The two-step system GMM estimator is more efficient than the one-step one. Nevertheless, it converges slowly to its asymptotic distribution and its standard errors tend to be biased downwards for finite samples. The one-step estimator does not experience such problem. The finite-sample correction to the two-step covariance matrix as derived by Windmeijer (2005) can be employed. The superiority of the two-step robust system GMM can be maintained. This approach uses multiple lags as instruments. The system is as a consequence over-identified. To test if the model is correctly specified and instruments are valid, both the Hansen (1982) J and the Sargan (1958) tests are computed. The latter is not robust to heteroscedasticity or autocorrelation, whereas, the former, which is the minimized value of the two-step GMM criterion function, is. The system GMM makes an exogeneity assumption where any correlation between endogenous variables and unobserved or fixed effects are constant over time. This allows the inclusion of levels equations in the system and use of lagged differences as instruments for these levels. The exogeneity assumption can be tested by using a difference-in-Hansen test (Bond et al. 2001).

Besides, the long-run impacts are required to test the aforementioned hypotheses. Bidirectional causality between LDEM_{it} and LGDP_{it} is found in the long-run. This situation is synonymous to endogenous regressors which can produce inconsistent and biased estimates. Long-run and efficient estimates can be obtained through the FMOLS and DOLS. These respective non-parametric and parametric methods can

Table 2 LLC panel unit root test statistics

Variables	Deterministics	Level form		First-difference	
		t value	t*	t value	t*
$LGDP_{it}$	Constant	−5.811	−1.50693 [0.066]‡	−12.581	−1.47332 [0.070]‡
	Constant + trend	−10.204	1.43561 [0.924]	−14.055	−0.23175 [0.408]
$LDEM_{it}$	Constant	−6.662	−0.15055 [0.440]	−16.121	−2.57481 [0.005]*
	Constant + trend	−10.637	0.60743 [0.728]	−17.537	0.11423 [0.5455]

Source computed. *Note* These LLC statistics are distributed as standard normal as both N and T grow large. Assuming no cross-country correlation and T is the same for all countries, the normalized t* test statistic is computed by using the t value statistics. After transformation by factors provided by LLC, the t* tests is distributed standard normal under the H_0 of non-stationarity. It is compared to the 1, 5, and 10 % significance levels with the one-sided critical values of −2.326, −1.645, and −1.282 correspondingly. The p values are in square brackets. *, **, and ‡ denote 1, 5, and 10 % significance level, respectively. These notations are applied for all succeeding tests

effectively correct for the biases resulting from the endogeneity of regressors over and above serially correlation and heteroskedasticity in error terms (Pedroni 2001). According to Kao and Chiang (2000), the DOLS method tends to outperform the FMOLS estimators in term of mean biases.

Equation (1), in conjunction with the long-run estimators, can be used to test the hypotheses linking democracy and economic growth. The Lipset hypothesis holds if, in the long-run, real GDP Granger-causes democracy and an increase in real GDP has a positive effect on democracy. Subsequently, the compatibility hypothesis holds if, in the long-run, democracy Granger-causes real per capita GDP and an increase in democracy results in an improvement in real per capita GDP. In contrast, the conflict hypothesis holds if, in the long-run, democracy Granger-causes real per capita GDP while a rise in democracy has a negative effect on real GDP. The skeptical hypothesis holds if there is no causal relationship between democracy and real GDP. Finally, the virtuous cycle hypothesis holds if bidirectionality prevails and each variable having a positive effect on each other.

4 Results

For implementation of the panel unit root tests, the Bartlett kernel is used. All bandwidths and lag lengths will be equal to $4(T/100)^{2/9} \approx 2.97$, where $T = 26$ (Basher and Westerlund 2008). The maximum lag length lies between 2 and 3. Too few lags may adversely affect the size of a unit root test, while too many lags can reduce the power of the unit root test (Campbell and Perron 1991). Martins (2011) applies 2 lags when performing the Pesaran panel unit root test over a period of 1980–2005. Since the time span is relatively short, the lowest value possible of 2 is chosen to conduct the panel unit root tests. The LLC test statistics are reported in Table 2. Both $LGDP_{it}$ and $LDEM_{it}$ are computed to be non-stationary and I(1), in keeping with the above-mentioned intuition about the order of integration.

The LLC panel unit root test assumes homogeneity in the AR(1) coefficients of the augmented Dickey–Fuller (ADF) specifications. Such assumption is quite implausible. Auxiliary tests, allowing for heterogeneity or cross-sectional dependence are needed to fully evaluate the order of integration of a series. The test assumes independently and identically distributed (iid) errors within their model. This assumption is violated in the presence of cross-sectional dependence. Residuals are hence contemporaneously correlated. In effect, cross-sectional dependence can lead to biased panel data unit root tests towards the alternative hypothesis (Banerjee et al. 2004). LLC (2002) recommends demeaning the series across N to attenuate this problem. Demeaned data refer to the extraction of the means from the time series. The rationale of demeaning is to remove the correlation in the sample. Such dependence can arise as a result of spatial spill-over effects, common unobserved shocks, social interactions, etc. (Breitung and Pesaran 2008).

The degree of cross-sectional dependence can be evaluated by calculating the pair-wise correlations[5] between changes in a variable (Koedijk et al. 2004). The pair-wise correlation coefficients of the first-differences in two series tend to be generally positive and quite large. For example, the correlation coefficient of ΔLGDP_{it} between Ivory Coast and Zimbabwe is equal to 0.4717 and for Central African Republic (Cen. Af. Rep.) and Liberia, it is 0.4878. The pair-wise correlation coefficients of ΔLGDP_{it} range from -0.6460 to 0.5869. Next, the pair-wise correlation coefficient of ΔLDEM between Congo Republic and Senegal is 0.3579 as well as for Burundi and Nigeria, it is 0.4221. Overall, the pair-wise correlation coefficients of ΔLDEM_{it} range from -0.7138 to 0.7508. These results show evidence[6] of a high degree of cross-sectional dependence within the SSA panel.

The IPS test allows for heterogeneity though it tends to have low power in panels with small T (Karlsson and Lothgren 2000). To some extent, it deals with cross-sectional dependence via demeaning. As revealed in Table 3, both $\text{LGDP}_{it} \sim\text{I}(1)$ and $\text{LDEM}_{it} \sim\text{I}(1)$. These tests are done using both the raw and demeaned data. Above and beyond, the ignorance of structural breaks can lead to a fall in power and to the rejection a unit root even although the trend stationarity alternative is true (Perron 1989). The ILT LM panel unit root test can account for such breaks and is regarded as being more powerful than the IPS test. Results in Table 4 demonstrate the rejection of the H_0 of a unit root after controlling for the presence one or two structural shifts in the trend.

It is not always simple to conclude about the order of integration of a series. Rejection of the H_0 of a unit root can be due to the existence of as few as one stationary series in the panel (Choi 2004). Kwiatkowski et al. (KPSS, 1992) recommend a test of the null of stationarity hypothesis to complement the null of a unit root one. This

[5] Detailed results of the pair-wise correlations are available upon request.

[6] For example, further evidence of cross-sectional dependence is obtained by estimating the following fixed-effects panel data model: $\text{LGDP}_{it} = \beta\text{LDEM}_{it} + u_{it}$, where u_{it} is the error term. When running the model, $\beta = 0.253$, with a p value of 0.000. The Pesaran (2004) test statistic of cross-sectional independence is equal to 37.369, with a p value of 0.000. The absolute value of the off-diagonal elements is next computed to be 0.548. These results reveal a high degree of cross-sectional dependence (De Hoyos and Sarafidis 2006). Moreover, groupwise heteroskedasticity is also found. Greene (1993) test statistics is equal to 2140.98 with a p value of 0.000. The null of homoskedasticity is rejected.

Table 3 IPS panel unit root test statistics

Variables	Data	Deterministics	Level form		First-difference	
			t-bar	Ψ_t	t-bar	Ψ_t
$LGDP_{it}$	Raw	Constant	−0.167	7.354 [1.000]	−2.837	−7.755 [0.000]*
		Constant + trend	−1.922	1.054 [0.854]	−3.034	−5.692 [0.000]*
	Demeaned	Constant	−1.523	−0.356 [0.361]	−2.408	−5.370 [0.000]*
		Constant + trend	−2.065	0.182 [0.572]	−2.562	−2.895 [0.002]*
$LDEM_{it}$	Raw	Constant	−1.336	0.707 [0.760]	−2.814	−7.627 [0.000]*
		Constant + trend	−1.867	1.385 [0.917]	−2.951	−5.202 [0.000]*
	Demeaned	Constant	−1.408	0.296 [0.616]	−2.840	−7.773 [0.000]*
		Constant + trend	−1.865	1.400 [0.919]	−3.049	−5.781 [0.000]*

Source Computed. *Note* t-bar is the panel test based on ADF statistics. Critical values for the t-bar statistics without trend at 1, 5, and 10 % significance levels are −1.820, −1.730, and −1.690 while with inclusion of a time trend, the critical values are −2.450, −2.370, and −2.320, respectively. Assuming no cross-country correlation and T is the same for all countries; the normalized Ψ_t test statistic is computed by using the t-bar statistics. The Ψ_t tests for H_0 of joint non-stationarity and is compared to the 1, 5, and 10 % significance levels with critical values of −2.330, −1.645, and −1.282 correspondingly. The p values are in square brackets

Table 4 ILT panel LM unit root test statistics

Variables	Level form	
	With one break	With two breaks
$LGDP_{it}$	−6.752*	−8.251*
$LDEM_{it}$	−7.644*	−10.267*

Source Computed. *Notes* Critical values for the LM panel unit root test are distributed asymptotic standard normal and are −2.326, −1.645, and −1.282 at the 1, 5, and 10 % levels, respectively. The minimum LM unit root test which accounts for a break in the trend is employed to test for the H_0 of non-stationarity

Table 5 Hadri panel unit root test statistics

Variables	Data	Deterministics	Level form	First-difference
			Z	Z
$LGDP_{it}$	Raw	Constant	20.3348*	1.1330
		Constant + trend	12.5066*	5.2247*
	Demeaned	Constant	17.9371*	1.0420
		Constant + trend	12.5911*	5.2570*
$LDEM_{it}$	Raw	Constant	16.2974*	−0.7847
		Constant + trend	8.5300*	2.3251**
	Demeaned	Constant	13.7646*	−0.7547
		Constant + trend	9.2276*	2.2868**

Source Computed. *Note* Hadri's test is based on the average of the N country-specific KPSS LM-statistics under which the H_0 of stationarity is tested. Heteroskedasticity is controlled while computing the statistics. Each Z statistic is compared to the 1, 5, and 10 % significance levels with the one-sided critical values of 2.326, 1.645, and 1.282, respectively

Table 6 Pesaran CADF panel unit root test statistics

Variables	Deterministics	Level form		First-difference	
		t-bar	Z	t-bar	Z
LGDP_{it}	Constant	−1.332	2.255 [0.988]	−2.480	−3.943 [0.000]*
	Constant + trend	−2.454	−0.817 [0.207]	−2.686	−2.142 [0.016]**
LDEM_{it}	Constant	−1.439	1.678 [0.953]	−2.635	−4.779 [0.000]*
	Constant + trend	−1.469	4.783 [1.000]	−3.003	−3.943 [0.000]*

Source Computed. *Note* The Pesaran CADF test of the H_0 of non-stationarity is based on the mean of individual DF (or ADF) t-statistics of each unit in the panel. Critical values for the t-bar statistics without and with trend at 1, 5, and 10 % significance levels are −2.300, −2.150, and −2.070; and −2.810, −2.660, and −2.580, respectively. Assuming cross-section dependence and T is the same for all countries. The normalized Z test statistic is computed by using the t-bar statistics. The Z test statistic is compared to the 1, 5, and 10 % significance levels with the one-sided critical values of −2.326, −1.645, and −1.282 correspondingly

can lead to a substantial gain in power of the testing framework. Such joint testing is commonly known as "*confirmatory analysis*" (Romero-Ávila 2008). The Hadri panel unit test offers a fitting alternative. These are based on the mean of KPSS test statistic. Table 5 shows the Hadri test statistics. The variables are found to be I(1).

One major drawback of the first generation of panel unit root tests lies in their assumption about cross-sectional independence. These tests tend to suffer from size distortions in the presence of cross-sectional dependence, leading to a low power of the testing framework (Herwartz and Siedenburg 2008). For instance, such dependence may have caused the failure to reject the alternative hypothesis of stationarity at level form for the ILT test. Most of the first generation tests resort to demeaned data to tackle this problem. This approach assumes the existence of one common factor with the same effect on all the units, which is quite restrictive. The demeaning of data may not "… *eliminate the size problem caused by the variation of cross correlations, and lead to false inference* (p. 309)" (Strauss and Yigit 2003).

The second-generation panel unit root tests explicitly control for cross-sectional dependence rather than using demeaned data. Pesaran (2007) suggests a test which allows for the presence of more general cross-sectional dependence patterns. To control for these patterns, the standard ADF regression models are augmented with the cross-section averages of lagged levels and first-differences of the individual series. The Pesaran test is based on the averages of the individual cross-sectionally augmented ADF (CADF) statistics. The test is found to have good size and power properties, even when N and T are relatively small. As presented in Table 6, results from the Pesaran test corroborate with the earlier tests. Both variables are once more found to follow an I(1) process.

A further challenge when carrying out a panel unit root test is to control for cross-sectional cointegration. Long-run dependence occurs when two or more units or countries share a common stochastic trend. This situation can again bias upwards the probability of Type I error of panel unit root tests whereby the H_0 is wrongly rejected (Banerjee et al. 2005). As such, cross-sectional cointegration can invalid not only first generation tests but also second-generation tests, such as the Pesaran test.

Table 7 Chang and Song panel unit root test statistics

Statistics	$LGDP_{it}$		$LDEM_{it}$	
	Level form	First-difference	Level form	First-difference
ta_c	2.08851	−7.39912*	−1.69250**	−12.36729*
ta_h	0.80839	−1.41420‡	−1.04353	−4.75161*
ta_a	0.75548	−0.41184	−0.94468	−2.62035*
tm_c	−1.52119	−2.91941‡	−2.21668	−5.33674*
tm_h	−1.08889	−1.84161	−1.54508	−5.48529*
tm_a	−0.61145	−2.07986	−1.12159	−3.33591**

Source Computed. *Note* The nonlinear IV average and minimum tests are denoted by the *ta* and *tm* while the subscripts c, h, and a refer to those tests with single IGF and no covariate, with single IGF and covariate and orthogonal IGF with no covariate, respectively. The tests include a constant term only. The H_0 of non-stationarity is tested. Each test statistic is compared to the 1, 5, and 10 % significance levels with the one-sided critical values of −2.326, −1.645, and −1.282 for the average test while these are −3.402, −2.928, and −2.696 for minimum test, respectively. The critical values for latter ($N = 30$) are computed by Chang and Song (2009)

Based on the Chang (2002) nonlinear IV panel unit root test, Chang and Song (2009) recommend a test which makes use of a set of orthogonal functions as instrument generating function (IGF) to tackle any forms of dependence. As exposed in Table 7, two different types of panel unit root tests are proposed. The average tests relate to the testing of the H_0 of non-stationarity for all individual units, whereas the minimum tests evaluate the H_0 of non-stationarity of some individual units in the panel. Three test statistics, such as ta_c, ta_h, and tm_c, confirm an I(1) process for $LGDP_{it}$, while all statistics, apart from the ta_c statistic, provide similar evidence for $LDEM_{it}$.

As a final and confirmatory test, the Hadri and Kurozumi test of the H_0 of stationarity is applied. This test is essentially an extension of the Hadri test and it allows the Lagrange multiplier (LM) test to control for cross-sectional dependence. Though it is similar to the KPSS test, the regression is augmented by cross-sectional average of the observations, in same way as the Pesaran test which augments the conventional ADF regression. As revealed in Table 8, two test statistics are computed. The ZA_spc and ZA_la are the augmented panel KPSS test statistics with long-run variance corrected by the Sul et al. (2005) and lag-augmented (Choi 1993; Toda and Yamamoto 1995) methods, respectively. $LGDP_{it}$ is found to be I(1) when referring to ZA_{la}. In contrast, $LDEM_{it}$ is computed to be I(0) as indicated by both statistics. Failure to reject the H_0 for $LDEM_{it}$ contradicts our expectation of non-stationarity. As indicated by Caner and Kilian (2001), unit root tests for the null of stationary, such as the KPSS test, tend to have serious size distortions when the H_0 is close to the alternative of a unit root. This situation may well be applied to panel unit root tests. In general, though there is overwhelming evidence supporting, an I(1) process both $LGDP_{it}$ and $LDEM_{it}$.

Various panel cointegration tests are next implemented. Table 9 reports the Nyblom and Harvey test statistics under both iid random-walk (RW) errors (NH-t) and the serially correlated residuals (NH adj-t) assumptions. No model needs to be estimated as the test is based on the rank of covariance matrix of the residuals driving the mul-

Table 8 Hadri and Kurozumi panel unit root test statistics

Statistics	Deterministics	$LGDP_{it}$ Level form	$LGDP_{it}$ First-difference	$LDEM_{it}$ Level form
ZA_{spc}	Constant	−1.28075	1.44300‡	0.09730
	Constant + trend	3.17061*	5.07164*	−1.07182
ZA_{la}	Constant	−1.77349	0.53785	0.13316
	Constant + trend	1.99839**	4.27586*	−0.96926

Source Computed. *Note* The H_0 of stationarity is tested. The ZA_{spc} and ZA_{la} test statistics is compared to the 1, 5, and 10 % significance levels with the one-sided critical values of 2.326, 1.645, and 1.282, respectively

Table 9 Nyblom–Harvey panel cointegration test statistics

Specifications	Statistics	$LGDP_{it}$	$LDEM_{it}$
Fixed effects	NH-t	12.9808*	12.4615*
	NH adj-t	134.4808*	133.3846*
Fixed effects and time trends	NH-t	11.1346*	11.5385*
	NH adj-t	88.4109*	114.8347*

Source Computed. *Note* The H_0 of the test is no cointegration (H_0 : rank(var–cov) $= K = 0$) against the alternative hypothesis of cointegration (H_1 : rank(var–cov) $= K \neq 0$). H_0: 0 common trends among the 28 series in the panel. NH-t: the test is performed under the hypothesis of iid errors. NH adj-t: errors are allowed to be serially correlated and the test is performed using an estimate of the long-run variance derived from the spectral density matrix at frequency zero. Critical values for the t-bar statistics without and with trend at 1, 5, and 10 % significance levels are 7.1862, 6.4117, and 6.0307; and 2.5905, 2.3997, and 2.3010, for N equals 30, respectively

Table 10 Pedroni panel cointegration test statistics

Statistics	Without trend	With trend
Panel ν	−1.74846	5.82008*
Panel ρ	0.77669	0.40157
Panel pp	0.23779	−1.44577‡
Panel adf	1.35077	−1.56545‡
Group ρ	2.80923	2.03288
Group pp	1.99732	−0.71100
Group adf	2.99250	−1.84306**

Source Computed. *Note* The H_0 of no cointegration is verified. These statistics are compared to a one-sided standard normal test with critical values of 1, 5, and 10 % given by −2.326, −1.645, and −1.282, respectively. A special case is the panel ν-statistic which diverges to positive infinity under the alternative hypothesis. The asymptotic distributions are derived in Pedroni (1999, 2004)

tivariate RW. Two specifications such as fixed effects without and with time trends are computed. The H_0 of no cointegration is rejected for both the iid RW NH-t and the non-parametric adjustment (with 2 lags) long-run variance, denoted by NH adj-t specification.

The results for the Pedroni test statistics are displayed in Table 10. Seven test statistics with the H_0 of no cointegration are computed. Four of these statistics, called

Table 11 Westerlund panel cointegration test statistics

Statistics	Without trend				With trend			
	Value	Z	p value	Robust p value	Value	Z	p value	Robust p value
Gt	0.203	5.999	1.000	1.000	−2.136	1.500	0.933	0.610
Ga	0.050	4.481	1.000	1.000	−8.415	2.799	0.997	0.344
Pt	1.021	3.189	0.999	0.878	−14.477	−3.772	0.000*	0.096‡
Pa	0.011	1.896	0.971	0.828	−6.800	1.840	0.967	0.344

Source Computed. *Note* The H_0 is no cointegration is tested. *Notes Gt*, *Ga*, *Pt*, and *Pa* are Westerlund (2007) cointegration test statistics in panel error-correction. All these statistics are distributed standard normally. Critical values of one-sided tests for 1, 5, and 10 % significance levels are −2.326, −1.645, and −1.282, respectively

panel cointegration statistics, are within-dimension-based statistics. These statistics are the panel ν, panel ρ, and panel pp, denoting the non-parametric variance ratio, Phillips–Perron ρ, and Student's t-statistics, respectively, whereas the panel adf is a parametric statistic based on ADF statistic. The other three statistics, called group mean panel cointegration statistics, are between-dimension-based statistics. These are the group ρ, group pp, and group adf, representing the Phillips–Perron ρ-statistic, Phillips–Perron t-statistic, and the ADF statistic, respectively. The three statistics allow for the modeling of an extra source of prospective heterogeneity across units. The H_0 cannot be rejected when the tests are executed when without a trend. On the contrary, when a time trend is included in the tests, H_0 is systematically rejected in four cases as confirmed by the panel ν, panel pp, panel adf, and group adf statistics. These two cointegration tests rely on the assumption of cross-sectional independence in the error term. This condition is unlikely to hold in practice.

The Westerlund test does control for cross-sectional dependence. Four test statistics are computed. *Ga* and *Gt* test statistics test the H_0 of no cointegration for at least one of the cross-sectional units. *Pa* and *Pt* test statistics use the pooled information over all the cross-sectional units to test the H_0 of no cointegration for the whole panel. The H_0 of no cointegration which infers whether the error-correction term in a conditional error-correction model is equal to zero is tested. If the H_0 of no error-correction is accepted, then the H_0 of no cointegration is also accepted. These tests have limiting normal distributions and are consistent. Results are reported in Table 11. To control for cross-sectional dependence, robust critical values is obtained through 500 bootstrap replications. The number of lags and leads are set to one. Similar to the Pedroni test, the H_0 cannot be rejected when a trend is excluded from the tests. But, the H_0 of no cointegration is rejected for the *Pt* test statistic when a trend is included in the test. The cointegration tests appear to be sensitive to the insertion of a time trend. Given the characteristics of the series, it is practical to integrate a deterministic time trend.[7]

[7] Hassler (2002) studies the implications of excluding a trend in time series cointegration tests. According to him, the asymptotic critical values of cointegration tests are affected by the presence of the linear trend in the regressors. Failure to account this fact can too often lead to tests which are biased towards establishing cointegration. Hence, this raises the need to consider a linear trend when testing for a long-run relationship.

Table 12 Blundell–Bond system GMM panel VAR causality test

Variables	$\Delta LGDP_{it}$	$\Delta LDEM_{it}$
$\Delta LGDP_{it-1}$	0.3364532	−0.7952365
	(0.1314181)**	(0.4563114)‡
$\Delta LDEM_{it-1}$	0.0336239	0.1810297
	(0.0667132)	(0.1343565)
ECT_{it-1}	−0.0059861	−0.1132842
	(0.0034986)‡	(0.0385936)*
Constant	0.0143265	0.0492928
	(0.0045967)*	(0.0120067)*
Observations	560	560
Number of instruments	28	28
Wald $\chi^2(3)$	8.14	20.68
	[0.043]**	[0.000]*
Sargan test of over-identifying restrictions	15.89	9.11
	[0.892]	[0.997]
Hansen test of over-identifying restrictions	25.32	23.22
	[0.388]	[0.507]
Difference-in-Hansen test of exogeneity	6.19	2.38
	[0.626]	[0.984]
AR(1) test of serial correlation	−2.01	−2.66
	[0.045]**	[0.008]*
AR(2) test of serial correlation	1.30	1.17
	[0.193]	[0.243]
Short-run causality test	0.25	3.04
	[0.614]	[0.081]‡
Long-run causality test	2.93	8.62
	[0.087]‡	[0.003]*

Source Computed. The model is estimated by the two-step system GMM. The robust standard errors are in parenthesis while p values are in square brackets. The explanatory variables are assumed to be endogenous and are instrumented in GMM-style *à la* Roodman (2006)

Table 12 presents the causality results of the Blundell–Bond system GMM panel VECM estimation. Due to rather small number of SSA countries in the sample, the use of too many instruments can cause the Sargan and Hansen tests to be weak. For small finite samples, a large number of weak instruments can overfull the endogenous variables and also reduce the accuracy in parameter estimations of the Sargan and Hansen tests of the instrument of joint validity (Roodman 2009). There is no formal test to detect "too many" instruments. One way to avoid such problem, is to use the rule of thumb of maintaining the number of instrument less than or equal to the number of groups (e.g., Docquier et al. 2011). As such, the number of instruments used is equal to 28.

Table 13 Panel FMOLS and DOLS estimates

Dependent variables	Independent variable							
	$LGDP_{it}$				$LDEM_{it}$			
	FMOLS		DOLS		FMOLS		DOLS	
	Coefficient	t-statistic	Coefficient	t-statistic	Coefficient	t-statistic	Coefficient	t-statistic
$LGDP_{it}$	–	–	–	–	0.30	3.65*	0.38	5.25*
$LDEM_{it}$	0.08	3.85*	0.08	5.00*	–	–	–	–

Source Computed. *Note* For the panel DOLS, maximum lag and lead length are set to 1 since $T < 30$ (Nelson and Donggyu 2003). For the FMOLS, the selection of bandwidth for kernels is automatically computed. The critical values of the two-tailed t-statistics test at 1, 5, and 10 % significance levels are 2.326, 1.645, and 1.282 for the panel, respectively

A telltale sign of valid instruments is a high p value of the Hansen J statistic of at least 0.25 (Roodman 2009). The Sargan and Hansen statistics are both greater than 0.25, confirming the validity of the instruments in use. Negative first-order serial correlation in the disturbances is discovered in the first differenced residuals. No second-order serial correlation is established. These results imply an absence of serial correlation among disturbances. Subsequent to the conjecture discussed previously, the lag order ρ of the panel VECM-based causality tests is computed to be 1. Real GDP is found to Granger-cause democracy in the short-run whereas a feedback relationship between real GDP and democracy prevails in the long-run. As observed in Table 12, the occurrence of endogenous regressors validates the use of efficient long-run estimators such as the FMOLS and DOLS. Autocorrelation is as well found when using the Wooldridge (2002) test. The autocorrelation statistic is computed as $F(1, 27) = 74.302$ with a p value of 0.000. To control cross-sectional dependence, common time dummies are included in the respective long-run estimators (Pedroni 2001). Referring to Table 13, both long-run estimators yield roughly similar results. The impact of democracy on economic growth and vice versa is found to be positive and is significant at 1 % level.

The above results reflect the importance of democracy in stimulating economic growth and vice versa. Narayan et al. (2011) find no systematic relationship between democracy and real GDP for the vast majority of countries (26 out of 30) when using the Freedom House index. His findings give support to the skeptical hypothesis for the SSA. Alternatively, by taking advantage of the increased power of the panel framework, such hypothesis is clearly rejected. In the short-run, democracy does not Granger-cause economic growth. This can be regarded as consistent with the skeptical hypothesis. But, in the long-run democracy and economic growth are found endogenously related. Democracy is found to Granger-cause economic growth and its impact on the latter is found to be positive. This is consistent with the compatibility hypothesis which displays the importance of democracy in promoting economic growth. Economic growth is also found to Granger-cause democracy in both the short-run and long-run and has a long-run positive impact. This is consistent with the Lipset hypothesis which highlights the significant role of economic growth to sustain democracy. All together, the virtuous cycle hypothesis is supported.

5 Conclusion and policy implications

The paper has examined the relationship between democracy and economic growth for a sample of 28 countries of SSA over the period 1980–2005. To conduct the study, a democracy index has been constructed from the Freedom House data. To make efficient inferences, various panel unit root and cointegration tests have been conducted. Both variables are found to follow an I(1) process. Evidence of a long-run relationship between democracy and economic growth is revealed. A panel VECM-based causality test is subsequently performed with the help of the Blundell–Bond two-step system GMM and the long-run impacts have been estimated using the panel FMOLS and DOLS estimators. Democracy is found to have no impact on economic growth in the short-run which is consistent with the skeptical hypothesis. Even so, economic growth does cause higher democracy in the short-run. In the long-run, democracy and economic growth are found to be mutually reinforcing with both having a positive impact on each other. This supports the virtuous cycle hypothesis.

These results have profound policy implications for the SSA. To some extent, the lack of democracy has gone a long way to explain the region's poor economic performance. Breaking this vicious cycle and turning it into a virtuous one is in the realm of the policymakers. Reforms in connection with the political system can be proposed without much concern about their impacts on economic growth in the short-run. Still, in the long-run, the strengthening of democracy is apt to improve economic growth whilst further economic growth can lead to the enhancement of democracy.

As implied by the results, free and fair elections and a reliable democratic system are important ingredients to ensure not only peace and political stability but are also part and parcel of a sound and smooth running economy (Sobhee 2009). Then again, democracy alone may not be sufficient. Adequate investment in human capital should equally take place. For instance, the deepening of education can improve the distribution of intellectual power resources and therefore reinforce the social basis of democratic political competition (Vanhanen 2004). Besides, democracy is intrinsically linked with financial development. Financial liberalization has long been considered as an engine of growth by the developed countries and has been prescribed to developing economies to enhance their catching-up process (Baltagi and Demetriades 2011). But, greater private liberty and freedom can lead to excessive financial deregulation and liberalization which in turn can cause financial crises (Lipscy 2011). An efficient financial system should also be regarded as the backbone of a healthy democracy. In sum, a democratic government can more efficiently deal with market failures such as limiting environmental degradation and catering for public and merit goods such as roads, defense, education, and health care assistance in a satisfactory manner. A well-functioning financial market can also be ensured by enacting proper laws and regulations.

So far, various efforts have been made by the SSA to support democracy and economic development. In September 2000, the SSA members, alongside other developing nations, have signed the Millennium Development Goals (MDGs) at United Nations Headquarters in New York, in a view to eradicate poverty and generate greater economic prosperity by 2015. The promotion and consolidation of democracy is a crucial item on the agenda. Along the same lines, the establishment of the African Union

(AU) in July 2002 was done with the main intention of promoting economic development across the African regions with special emphasis being laid on democratic institutions, good governance and human rights. The establishment of the 2001 New Partnership for Africa's Development (NEPAD) programme under the aegis of the African Union illustrates the will of the African states to eliminate conflicts and push towards a democratic system of governance. Its objectives, which are close to those of the MDGs, are aimed at fostering Africa's growth, development, and participation in the global economy. However, to be able to operate in an effective way, the credibility of the AU and NEPAD should be maintained by holding each member state to adhere to high standards of democracy and imposing consequences for those states not abiding to them. To sum up, democracy is the key to unlocking the door of greater economic prosperity in the SSA.

Acknowledgments I would express my appreciation to Thomas F. Rutherford, Ron Schoenberg, Denise Stark, Badi H. Baltagi and two anonymous referees for their comments and suggestions. My thanks are also extended to Junsoo Lee, Yoosoon Chang, Wonho Song, Kaddour Hadri, and Eiji Kurozumi for sharing their GAUSS codes. Errors, if any, are the author's own solely.

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