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International remittances, poverty and growth into WAEMU countries: evidence from panel cointegration approach

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Abstract

This paper applies the panel unit root, panel cointegration and panel vector error-correction Granger causality tests in a panel of the eight West African Economic and Monetary Union (WAEMU) countries and shows that in the short-run, there is evidence of (i) a bidirectional causal relationship between economic growth and poverty; (ii) a unidirectional causal relationship from remittances to poverty; and (iii) a unidirectional causal relationship from economic growth to remittances. Our results further suggest that in the long-run, both economic growth and remittances jointly Granger-cause poverty and that it takes more than fifteen years for poverty to converge to its long-run equilibrium in response to changes in economic growth and remittances.

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1 Introduction

Remittances inflows have increased substantially over the last three decades and have become the main financial external inflow in some developing countries, surpassing other inflows that traditionally play an important role in these countries, such as official development assistance (ODA) and foreign direct investment (FDI). According to World Bank, official remmittances[1] grew by 9.6 per cent in 2018 (up from the 8.8 per cent rise in 2017), to reach a record $529 billion [9]. The trend is particularly true of Sub-Saharan Africa (SSA). Remittances flows to the region have increased steadily over the last three decades from about 0.5 per cent of regional GDP in 1980 to almost 3 per cent in 2018 [8], with five of the top 25 countries with the greatest remittance share of GDP in 2009 being located in this region.[2] Moreover, SSA countries have received $45 billion in remittances in 2018 and on average, the region has perceived 7% of world remittances inflows over the period 2010-2018 [23].
Like for many SSA sub-regions, remittances have become an increasingly important source of financing for the members of the West African Economic and Monetary Union (WAEMU).[3] According to the World Bank, the inflows of remittances in WAEMU countries has quadrupled between 2000 and 2011.[4] Given this huge increase, it is not surprising that remittances have a positive impact on growth and poverty reduction in the WAEMU. Indeed, the bulk of remittances in the union (54 per cent), that are in many cases directly received by the poor, is spent on immediate consumption necessities such as food, clothing, medicine and shelter. Hence, remittances help lift huge numbers of people out of poverty by supporting a higher level of consumption than would otherwise be possible [21, 16, 33]. This increased consumption will generate multiplier effects in the economy, which in turn will have a positive impact on growth[5] [2, 35]. In addition, remittances within the WAEMU are countercyclical and play an important role as shock-absorbing device when economies slow down.[6]
The past few years have witnessed a remarkable revival of interest in the impact of international remittances in recipient country, notably on economic growth and poverty reduction. More specifically, the bulk of previous studies investigated the relationship between remittances and economic growth and failed to reach a consensus as to the direction of causation. With the exceptions of Siddique et al. [40] and Donou-Adonsou and Lim [12] who concluded that in India and in the WAEMU countries respectively, remittances and economic growth were independent, the balance of the literature finds that causality either runs from remittances to growth (Jawaid and Raza [20]; Olubiyi [30]); from growth to remittances (Ali et al. [4]), or that there is a bidirectional relationship between remittances and growth (Kumar and Vu [24]; Ahmed and Hakim [3]).
However, there is far less evidence on the causal relationship between remittances and poverty reduction and the findings of the few studies on the topic are inconclusive. Gaaliche and Gaaliche [15] and Hatemi and Uddin [18] found a bidirectional causality relationship between remittances and poverty reduction while Muhammad et al. [6] uncovered a unidirectional causality running from remittances to poverty. Furthermore, studies that examined the long-run or the short-run impact of remittances at country level focused mostly on a few sample of emerging or high-income countries, leaving SSA countries with no coverage (notable exceptions include Olubiyi [30], Ahmed and Hakim [3] and Donou-Adonsou and Lim [12]).
In this context, this study contributes to the existing literature by examining the short-run and the long-run relationships between remittances, economic growth and poverty reduction concurrently, on a data set of the eight WAEMU countries from 2000 to 2018 using panel unit root, panel cointegration and panel vector error-correction Granger causality tests. WAEMU countries are chosen mainly because over the past two decades, like other SSA regions, they have experienced growing remittance inflows, on the one hand, and high levels of poverty as well as consistently low levels of economic growth, on the other hand [34]. Better understanding the nexus between remittances, growth and poverty reduction could help policy makers and financial institutions to design appropriate policy instruments to maximize the poverty reduction and developmental impacts of remittance flows.
The rest of the paper proceeds as follow: the second section describes the empirical model and the data. Section three follows up with the econometric techniques while section four discusses the empirical results. A final section concludes and provide some policy recommendations.

[1]Freund & Spatafora [14] estimate that the amount of informal remittances sent through informal channels, e.g. self-carry, hand-carry by friends or family members or in-kind remittances of clothes and other consumer goods may equal about 35 to 75 per cent of official flows.

[2]Lesotho, Togo, Cape Verde, Guinea-Bissau and Senegal.

[3]Members of the WAEMU (also known by its French acronym, UEMOA) are Benin, Burkina Faso, Côte D'Ivoire, Guinea-Bissau, Mali, Niger, Senegal, and Togo.

[4]While part of this increase is likely due to better measurement, it also reflects higher migration from Africa and higher incomes of African migrants [33].

[5]Ratha[35] suggests that increased remittances might have substantial multiplier effects, because they are more likely to be spent on domestically produced goods.

[6]International Monetary Fund (2013), Country Report 13/92. West African Economic and Monetary Union (WAEMU): Staff Report on Common Policies for Member Countries.

2 Empirical model and data

In this section, we adopt an empirical model that captures the relationship between remittances, economic growth and poverty reduction, and describe the data.

2.1 Empirical model

To investigate the causal relationship between remittances, growth and poverty reduction, we use the extended version of the basic growth-poverty model suggested by Ravallion and Chen [36], in which poverty can be modeled as a function of mean income, some measure of income distribution, and a variable of interest. In doing so, we follow Adams and Page [1], Gupta et al. [17]. The relationship that we want to estimate can therefore be expressed as follows:

$$ln(pov)_{it} = \alpha_i + \beta_1 ln(gdp)_{it} + \beta_2 ln(rem)_{it} + \epsilon_{it} \tag{1}$$

Where $i = 1, 2, ...N$ and $t = 1, 2, ...T$ are country and time notations, $(pov)_{it}$ is the measure of poverty in country i at year t, ranging from 0 to 1, economic growth is represented by $(gdp)_{it}$ and $(rem)_{it}$ is the international remittance flows as a percentage of GDP. The β coefficients in (1) capture the long-run effects between the variables, while α_i are country specific fixed effects that help to control any omitted factors that are stable over time. All time-varying variables are expressed in natural logarithms.

2.2 Data and descriptive statistics

Our empirical analysis is based on a balanced panel for the eight WAEMU countries covering the period 2000-2018. What these countries share in common is that most of them are among the poorest and least developed countries in the world. They have also experienced a major increase in remittance inflows over the past decades. The data used in this paper was mainly collected from the World Bank's World Development Indicators (WDI). The poverty headcount ratio is the percentage of people living on less than $1.90 a day. Remittances[7] are expressed as a ratio of GDP of recipient countries. The economic growth variable is the per capita GDP in constant 2010 United States dollars.
Table 1 reports the sample summary statistics. For all variables, we observe a significant difference between the maximum and the minimum values over the 18 years. This can be explained by the fact that our variables either constantly increased (remittances and GDP) or decreased (poverty headcount). Indeed, the high poverty levels in sources countries as well as the widening income inequalities between source and destination countries over the past decades led more people to choose to migrate. At the same time, remittances transfer costs have been lowered, so that more remittances could be sent. Increased remittances then raised standard of living of the poor and led to a reduction of their poverty headcounts.

3 Cointegration analysis

Following the existing literature, we start our analysis by examining stationarity of our variables using several first generation tests. We then apply the seven panel cointegration tests proposed by Pedroni [31] to obtain the long-term relationship between all variables. This is followed by the estimation of the cointegration coefficients through Panel Dynamic Ordinary Least Squares (DOLS) and Panel Fully Modified Ordinary Least Squares (FMOLS) estimators. Finally, Panel Vector Error Correction based Granger causality test is used in order to examine the causal relationship between variables.

3.1 Panel unit root test

Prior to using panel cointegration approach to investigate the possibility of long-run cointegration among the variables, it is necessary to determine the existence of unit roots in data series. There are several panel unit root tests in the literature, but the ones proposes by Levine, Lin and Chu [25](LLC), Im, Pesaran, and Shin [19](IPS), Maddala and Wu [26](Fisher-ADP and Fisher-PP) and Breitung [10] are used in this paper.
The LLC test which is the most widely used panel unit root test is based on the following Augmented Dickey Fuller (ADF) regression, but in panel settings:

[7] In this work as in most studies on the topic, remittances are aggregate worker remittances, compensation to employees, and migrant transfers series from the IMF Balance of Payments database, supplemented by the data from World Bank.

$$\Delta y_{it} = \alpha y_{it-1} + \sum_{j=1}^{p_i} \beta_{ij} \Delta y_{it-j} + \delta_i Z_{it} + u_{it} \tag{2}$$

Where p_i is the lag length, Z_{it} is a vector of deterministic terms, explaining the fixed effects or individual trends, and δ_i is the corresponding vector of coefficients.
Since the lag length of the differenced terms (p_i) is unknown, Levin et al [25] suggest the following three-step procedure: (i) carry out separate ADF regressions for each individual and generate two orthogonalised residuals; (ii) estimate the ratio of long-run to short-run innovation standard deviation for each individual; (iii) compute the pooled t-statistics, with the average number of observations per individual and average lag length. In this test, the associated autoregressive coefficient is constrained to be homogenous across individuals (i.e. $\alpha_i = \alpha$ for all i). Hence, the null hypothesis assumes a common unit root ($H_0 : \alpha = \rho - 1 = 0$) against the alternative hypothesis that each time series is stationary ($H_1 : \alpha < 0$).
Maddala [26] pointed out that because ρ is homogeneous across all i, the null may be fine for testing convergence in growth among countries, but the alternative restricts every country to converge at the same rate. This drawback led Im et al [19] to extend the LLC test by allowing heterogeneity on the autoregressive coefficient. In practice, their test takes the average of the individual ADF statistics across sections and standardizes it with the expected mean and variance. This approach allows for different specifications of the coefficients (α_i for each cross-section), the residual variance and lag-length [5]. The authors propose the following $t - bar$ statistic, based on the average of the individual unit root (ADF) test statistics:

$$\bar{t}_{IPS} = \frac{\sqrt{N}[\frac{1}{N}\sum_{i=1}^{N} t_i - \frac{1}{N}\sum_{i=1}^{N} E(t_i|\rho_i = 0)]}{\sqrt{\frac{1}{N}\sum_{i=1}^{N} var(t_i|\rho_i = 0)}} \Longrightarrow N(0,1) \tag{3}$$

This statistic evaluates whether the coefficient α is non-stationary across all individuals ($H_0 : \alpha_i = 0$ for all i), against the alternative hypothesis that at least a fraction of the series is stationary ($H_1 : \alpha_i < 0$ for at least one i).
Using Monte Carlo simulation, Breitung [10] shows that the power of the LLC and IPS tests statistics is sensitive to the specification of the deterministic components, such as the inclusion of individual specific trends [7]. He proposes a test statistic based on modifications to the LLC steps to overcome these difficulties. The null hypothesis of Breitung's test is that the panel series exhibits non-stationary difference, and the alternative hypothesis assumes that the panel series is stationary.
Maddala and Wu [26] and Choi [11] suggest the use of non-parametric Fisher tests. The main feature of these tests is that they combine the probability limit values ($p - values$) of unit root tests from each cross-section rather than average test statistics. Fisher tests are usually implemented using individual ADF or Phillips-Perron unit root tests, and their asymptotic distribution follows a chi-square ($P - test$). The test statistic is expressed as follows:

$$P = -2\sum_{i=1}^{N} ln\beta_i \Longrightarrow \chi^2_{2N} \tag{4}$$

The null hypothesis is that each series in the panel has a unit root, i.e., $H_0 : \rho i = 0$ for all i and the alternative hypothesis is that not all of the individual series has a unit root, i.e., $H_1 : \rho i < 0$ for $i = 1, .., N_1$ and $\rho i = 0$ for $i = N_1 + 1, ...N$.

3.2 Pedroni's panel cointegration tests

After testing for stationarity of the variables, the next step is to check for the existence of a long-run relationship among them. For that purpose, we applied the residual-based method developed by Pedroni [31]. Pedroni [31] proposes cointegration tests for heterogeneous panels based on the two-step cointegration approach of Engle and Granger [13]. Pedroni uses the residuals from the static (long-run) regression and constructs seven statistics to test the null hypothesis of no cointegration against the alternative hypothesis of cointegration. Four of them are based on pooling (within-dimension or 'panel statistics test'), which assumes homogeneity of the autoregressive term, whilst the remaining are less restrictive (between-dimension or 'group statistics test') as they allow for heterogeneity of the autoregressive term. In the case of panel statistics, ρ is assumed to be the same across all the cross sections, thus the null and alternative hypotheses are $1 :\ H_0 : \rho_i = 1$ for all i and $H_1 : \rho_i = \rho < 1$, respectively. In the case of group panel statistics ρ is allowed to vary over the cross sections, thus the null and alternative hypotheses are $H_0 : \rho_i = 1$ for all i and $H_1 : \rho_i < 1$ for at least one i .

The starting point for the cointegration tests for heterogeneous panels proposed by Pedroni [31] is the estimation of the following panel regression:

$$ln(pov)_{it} = \alpha_{it} + \delta_{1i} ln(gdp)_{it} + \delta_{2i} ln(rem)_{it} + \epsilon_{it} \tag{5}$$

for $i = 1, ..., N,\ t = 1, ..., T$; where T refers to the number of observations over time and N refers to the number of individual members in the panel. Pedroni [31] proposes the following steps: First, estimate the panel cointegration regression, include any desired intercepts, time trends, or common time dummies in the regression and collect the residuals $\hat{\epsilon}_{it}$ for later use. Second, differentiate the original data series for each member and compute the residuals for the following differentiated regression:

$$\Delta ln(pov)_{it} = \delta 1i \Delta ln(gdp)_{it} + \delta_{2i} \Delta ln(rem)_{it} + \eta_{it} \tag{6}$$

Third, calculate $\widehat{L}^2_{11i}$ as the long-run variance of $\widehat{\eta_{it}}$. Fourth, use residual $\widehat{\epsilon}_{it}$ of the original cointegrating equation and estimate the appropriate autoregressive model. To compute the nonparametric statistics, estimate $\widehat{\epsilon}_{it} = \widehat{\psi}_i \widehat{\epsilon}_{it-1} + \widehat{u}_{it}$, and use the residuals to compute the long-run variance of $\widehat{u}_{it}$, denoted $\widehat{\sigma}^2_i$. The term $\widehat{\lambda}_i$ can then be computed as $\widehat{\lambda}_i = \frac{1}{2}(\widehat{\sigma}^2_i - \widehat{s}^2_i)$ where $\widehat{s}^2_i$ is the simple variance of $\widehat{u}_{it}$. For the parametric statistics, estimate $\widehat{\epsilon}_{it} = \widehat{\psi}_i \widehat{\epsilon}_{it-1} + \sum_{k=1}^{k_i} \widehat{\psi}_{ik} \Delta \widehat{\epsilon}_{i,t-k} + \widehat{u}_{u,t}$. Using each of these parts, construct the seven statistics (see appendix) and then apply the appropriate mean and variance adjustment terms reported in Pedroni [31]. The asymptotic distributions for each of the seven panel and group mean statistics can be expressed in the form:

$$\frac{X_{N,T} - \mu\sqrt{N}}{\sqrt{\nu}} \Longrightarrow N(0,1)$$

where $X_{N,T}$ is the appropriately standardized form for each of the seven statistics, and the values for μ and ν are functions of the moments of the underlying Brownian motion function. The panel V-statistic is a one-sided test where large positive values reject the null of no cointegration. The remaining statistics diverge to negative infinitely, which means that large negative values reject the null. The critical values are also tabulated by Pedroni [31].

3.3 Panel cointegration coefficients estimation

In the presence of cointegrating variables, one would like to be able to efficiently estimate and test the relevant cointegrating vectors. Since our variables are cointegrated, and thus a long-run relationship exists among them, we estimate equation (1) by employing DOLS and FMOLS[8] introduced by Pedroni [32]. We opt for this estimator since it yields unbiased and asymptotically efficient estimates of the long-run relationship, even if there are endogenous regressors, thus allowing to control for the potential endogeneity of remittances, growth and poverty. It is well known that Ordinary Least Squares (OLS) estimation yields biased results because the regressors are endogenously determined in the case of variables which are integrated in order one or I(1).
To present the method, we consider the following fixed-effect panel cointegration system:

$$y_{it} = \alpha_i + x\prime_{it}\beta \tag{7}$$

where $t = 1, ..., T$ and $i = 1, ..., N$
$x\prime_{it}$ can in general be m-dimensional vectors of regressors which are integrated of order one, that is:

$$x_{it} = x_{i,t-1} + u_{2,it} \forall i, t \tag{8}$$

where the vector error process $w_{it} = (u_{1,it}; u_{2,it})$ is stationary with the asymptotic covariance matrix Ω_i, $\forall i = 1, ..., N$, $\Omega_i = \Omega^0_i + \Gamma_i + \Gamma\prime_i$, Ω^0_i is the contemporaneous covariance and Γ_i is the weighted sum of autocovariances. The long-run covariance matrix is constructed as follows:
$\begin{bmatrix} \Omega_{11i} & \Omega\prime_{21i} \\ \Omega_{21i} & \Omega_{22i} \end{bmatrix}$, where Ω_{11i} is the scalar long-run variance of the residual ϵ_{it}, Ω_{22i} is the long-run covariance among the $u_{2,it}$, and Ω_{21i} is the vector that gives the long-run covariance between the residual $u_{1,it}$ and each of the $u_{2,it}$
The FMOLS estimator is given by:

$$\widehat{\beta}_{FMOLS} = \left(\sum_{i=1}^{N} \widehat{L}^{-2}_{22i} \sum_{t=1}^{T} \left(x_{it} - \bar{x}_i \right)^2 \right)^{-1} \sum_{i=1}^{N} \widehat{L}^{-1}_{11i} \widehat{L}^{-1}_{22i} \left(\sum_{t=1}^{T} \left(x_{it} - \bar{x}_i \right) y^{\star}_{it} - T\widehat{\gamma}_i \right) \tag{9}$$

[8] Kao and Chiang [22] used Monte Carlo simulation to show that DOLS outperforms both the OLS and the FMOLS even though the coefficients of the DOLS converge to the same coefficients as the FMOLS estimation.

where

$$y_{it}^{\star} = (y_{it} - \bar{y}_i) - \frac{\widehat{L}_{21i}}{\widehat{L}_{22i}} \Delta x_{it} + \frac{\widehat{L}_{11i} - \widehat{L}_{22i}}{\widehat{L}_{22i}} \beta (x_{it} - \bar{x}_i)$$

and

$$\widehat{\gamma}_i = \widehat{\Gamma}_{21i} + \widehat{\Omega}_{21i}^0 - \frac{\widehat{L}_{21i}}{\widehat{L}_{22i}} \left(\widehat{\Gamma}_{221} + \widehat{\Omega}_{21i}^0 \right)$$

The panel group FMOLS estimator is the average of the FMOLS estimator computed for each individual:

$$\widehat{\beta}_{FMOLSG} = N^{-1} \sum_{i=1}^{N} \widehat{\beta}_{FMOLS} \tag{10}$$

3.4 Panel granger causality tests

Once the variables are cointegrated, we have to examine the direction of causality between them. Engle and Granger [13] show that if two non-stationary variables are cointegrated, a vector autoregression (VAR) in first differences will be misspecified. To test for Granger causality therefore, we use a vector error correction model (VECM). This means that the traditional VAR model is augmented with a one-period lagged error correction term that is obtained from the cointegrated model based on OLS. The VEC term represents any deviation from the long-run equilibrium between the dependent and the independent variables. The Granger causality test is based on the following regressions:

$$\Delta pov_{it} = \alpha_{1i} + \sum_{k=1}^{q} \theta_{11ik} \Delta pov_{it-k} + \sum_{k=1}^{q} \theta_{12ik} \Delta gdp_{it-k} + \sum_{k=1}^{q} \theta_{13ik} \Delta rem_{it-k} + \lambda_1 EC_{it-1} + \mu_{1it} \tag{11}$$

$$\Delta gdp_{it} = \alpha_{2i} + \sum_{k=1}^{q} \theta_{21ik} \Delta gdp_{it-k} + \sum_{k=1}^{q} \theta_{22ik} \Delta pov_{it-k} + \sum_{k=1}^{q} \theta_{23ik} \Delta rem_{it-k} + \lambda_2 EC_{it-1} + \mu_{2it} \tag{12}$$

$$\Delta rem_{it} = \alpha_{3i} + \sum_{k=1}^{q} \theta_{31ik} \Delta rem_{it-k} + \sum_{k=1}^{q} \theta_{32ik} \Delta pov_{it-k} + \sum_{k=1}^{q} \theta_{33ik} \Delta gdp_{it-k} + \lambda_3 EC_{it-1} + \mu_{3it} \tag{13}$$

Here Δ is the first difference operator, α_i represent country-specific fixed effects, θ_i the coefficients corresponding to the k^{th} lag of the endogenous variables; λ_i the coefficients of the error correction terms and μ_k the idiosyncratic errors. In addition to the variables defined above, ECT_{t-1} is the one period lagged error-correction term derived from the cointegrating equation. A VECM allows testing for both short and long-run causality. In the system of equations (11)-(13), the coefficients θ_i represent the short-run effect of the endogenous variables. A standard Wald test on these coefficients can be used to test for short-run Granger causality. Specifically, we test the null hypothesis $H_0 : \theta_i = 0$. In equation (11), rejecting the null hypothesis $\theta_{12} = 0$ implies that gdp Granger causes change in poverty in the short-run. In other words, the first lag of gdp is a significant predictor of changes in poverty. Similarly, rejecting the null hypothesis $H_0 : \theta_{22} = 0$ in equation (12) implies that gdp is responding to short term shocks in poverty. The joint significance of the coefficients θ_{12} and θ_{22} implies bidirectional causality in which the two variables Granger-cause each other in the short run, while the rejection of only one of the hypotheses implies unidirectional causality.

We can test for long-run Granger causality between variables in our model by examining the significance of the coefficient λ_k, which represents the speed of adjustment to the long-run equilibrium in response to any shocks to the system. We test the null hypothesis $H_0 : \lambda_k = 0$. The rejection of the null hypothesis implies long-run Granger causality running from the independent variables to the respective dependent variable. For example, the significance of λ_1 in equation (11) implies that changes in poverty adjust in the long-run to any temporary deviations from economic growth and remittance flows.

4 Empirical results

We started our empirical investigations using several first generation unit root tests. Given that we are dealing with macroeconomic variables that are often found to be non-stationary [28] and thus conducive to spurious regression, testing for stationarity is a crucial step prior to undertaking panel cointegration analysis. Once the variables are found to be integrated of the same order, we test for the long-term relationship between them. The cointegration coefficients are then assessed through DOLS and FMOLS estimators. Finally a panel VECM is estimated to evaluate the short and long-run impacts.

4.1 Panel unit root results

Table 2 reports the results of the panel unit root tests for the level and first differenced series of remittances, GDP and poverty. In level form, all the tests cannot reject the null hypothesis of unit roots, except for remittances. However, after applying the first difference, the null hypothesis is rejected for all tests at the 1% level. Overall, all variables are found to be integrated of order one, or I (1).

4.2 Panel cointegration results

Having found that the three variables are integrated of order one, the next step before testing Granger causality is to conduct cointegration tests. Results in table 3 show that four out of the seven Pedroni [31] statistics reject the null hypothesis of no cointegration. Hence, while unit root tests provided support for the presence of stochastic trends in the data, cointegration tests suggest that these trends have cancelled each other out – leading to stationary residuals. In practice, this means that these variables have a significant long-run relationship.

4.3 Granger causality test results

Based on the panel cointegration test results, we know that there is the presence of a long-run relationship between variables. However, the cointegration test results don't give information about the direction of this relationship. Granger causality test results are reported in table 4 below.

The short-run causality tests are performed through the Wald χ statistics, while long-run causality is inferred from the coefficients of ECT and corresponding t-statistics. In the short-run, there is evidence of (i) a bidirectional causal relationship between economic growth and poverty ($gdp \Leftrightarrow pov$), (ii) a unidirectional causal relationship from remittances to poverty ($rem \Rightarrow pov$) and (iii) a unidirectional causal relationship from economic growth to remittances ($gdp \Rightarrow rem$).

Our short-run results can be contrasted with Nyasha et al. [29] who found that there is a two-way relationship between economic growth and poverty reduction in the case of Ethiopia. Muhammad et al.[6] found a unidirectional causality running from remittances to poverty in lower-middle and upper-middle income countries while Ali et al. [4] uncovered a unidirectional causal relationship from economic growth to remittances. The short-run causality from growth to remittances is also consistent with findings of the IMF[9], who found that a one per cent decrease in real WAEMU GDP is associated with a four per cent increase in the net remittances inflows to the WAEMU (about 0.1 per cent of GDP). This means that migrants tend to send more money home when economic activity in a recipient country slows down, and vice versa, which confirms the countercyclical character of remittances.

For the long-run causality results, only the coefficient of the ECT when poverty is the dependent variable, is negative and statistically significant. This implies that (i) poverty tends to converge to its long-run equilibrium in response to changes in growth and remittances, and (ii) both growth and remittances jointly Granger cause poverty reduction in the long-run (gdp & $rem \Rightarrow pov$). It should also be noted that the ECT coefficient of 0.06 means that it takes more than fifteen years (1/0.06) for poverty to return to equilibrium after a shock. The ECT coefficient also means that about six percent of this disequilibrium is corrected in 1 year. In contrast, there is no evidence of a long-run relationship or causality when remittances and growth are dependent variables.

Both the short- and long-run Granger causality confirm that causality runs from economic growth and remittances to poverty. The short-run causality from remittances to poverty suggests that remittances are a short-run strategy for migrants originating from WAEMU countries to help their family left behind to escape poverty. The long-run joint causality for remittances and economic growth to poverty suggests that WAEMU countries that receive remittances generally face high levels of poverty and low levels of economic development. The lack of a long-run relationship when GDP is the dependent variable implies that remittances cannot be a long-term solution for growth within the WAEMU.

After estimation of the panel VECM equations, it is important to perform panel data serial correlation tests to confirm the validity of the panel VECM estimations [42]. For that purpose, we used the Breusch-Godfrey serial correlation Test. The null hypothesis is that there is no serial correlation against the alternative that there is serial correlation. Results show that all three equations do not have serial correlations. The $p - values$ for all three equations are more than 10%, and therefore we cannot reject the null hypothesis, which means that all the equations are free from serial correlation.

[9] International Monetary Fund (2013), Country Report 13/92. West African Economic and Monetary Union (WAEMU): Staff Report on Common Policies for Member Countries.

4.4 Long-run relationship coefficients estimation

Having established cointegration as well as the direction of causality in the short and in the long-run, we examine the long-run elasticities of the impact of remittances and economic growth on poverty reduction. The two long-run estimators that we use for this purpose are the FMOLS and DOLS. Both FMOLS and DOLS display similar results, in terms of the sign and statistical significance - remittances and growth have a negative effect on poverty - , whereas the magnitudes of the estimated coefficients are slightly different. Because the variables are expressed in natural logs, the coefficients on the remittances and growth variables can be interpreted as elasticities. All the coefficients are statistically significant at the 1% level of significance. The results suggest that a one percent increase in growth leads to a reduction of the poverty headcount ratio by 0.72-0.78 percent while the same increase in remittances decreases poverty by 0.13-0.15 percent. The elasticity for remittances with respect to poverty is close to Tsaurai's [41] findings for panel data of selected emerging countries, which he found to be equal to be 0.16.

5 Conclusion

This paper examines the causal relationship between international remittances, economic growth and poverty reduction in a panel of the eight WAEMU countries using panel unit root, panel cointegration and panel vector error-correction Granger causality tests. The main findings are that in the short-run, there is evidence of (i) a bidirectional causal relationship between economic growth and poverty; (ii) a unidirectional causal relationship from remittances to poverty; and (iii) a unidirectional causal relationship running from economic growth to remittances. In the long-run, we find that (i) it takes more than fifteen years to poverty to converge to its long-run equilibrium in response to changes in remittances and economic growth and (ii) remittances and economic growth jointly Granger cause poverty reduction.

The findings on the short-run causality can be contrasted with those of the IMF [10]which found that remittances within the WAEMU play an important role as a shock-absorbing device when economies slow down.

Our long-run causality results suggest that the WAEMU, like the bulk of developing countries, partly rely on remittances to reduce their levels of poverty and foster their economic growth. This is line with United Nations 2030 Agenda for Sustainable Development where they recognize that migration and remittances could contribute to the long-run development of the receiving countries [27].

To facilitate remittances flows and take full advantage of their economic development and poverty reduction potential, some measures have to be taken: (a) reduce barriers to migration such as legal restrictions and costs incurred so that more people will be able to migrate legally and send remittances home; (b) increase transparency and competition in the transfer market with the aim of reducing the cost of sending money home. This measure will not only make more money to flow directly into the pocket of the recipients, it will also lower the amount of remittances that go unrecorded because of high transaction fees; (c) increase banking penetration in the WAEMU which remains low. The ratio of adult with a bank account is about 36 per cent [39], significantly below the SSA average of 43 per cent. Higher access to financial institutions would channel more transfer into the formal sector. Furthermore, mobile banking, which has been a significant source of domestic transfer in east Africa, remains little developed in the western part of the continent. The development of mobile banking in the WAEMU could facilitate substantially intra-waemu remittances flows, as there would be no exchange rate cost associated with money transfers; (d) take measures to ensure that remittances recipients have access to targeted financial services to help them save and/or invest their funds and access credit.

However, the results from this paper should be interpreted with caution given the relatively short time-span of the data and the type of poverty measure used. Further work will be needed to refine the analysis as new, and improved data become available that spans for much longer period of time than in this study. Furthermore, the use of non-income poverty measures will give a more realistic picture of the state of poverty. Indeed, even though income helps to smooth the consumption pressure and reduce poverty, other aspects of life deserve attention. In measuring poverty therefore, it is necessary to go beyond income to incorporate other dimensions that are valuable to people such as basic education, basic health care or access to safe drinking water or basic sanitation [38]. In this framework, someone who has income and lacks access to basic education, basic health care or safe drinking water should be deemed poor. A study that will assess the interaction between remittances, growth and non-income poverty will be a good complement to this study.

[10]International Monetary Fund (2013), Country Report 13/92. West African Economic and Monetary Union (WAEMU): Staff Report on Common Policies for Member Countries.

Tables and Pedroni statistics

Tables

Table 1: Descriptive Statistics

Variables	Poverty headcount	GDP per capita	Remittances
Mean	48.02	760.48	4.20
Median	49.7	657.39	3.16
Maximum	74.9	1692.54	10.70
Minimum	22.9	322.78	0.80
Std. dev.	12.09	341.25	2.92

Table 2: Panel Unit Root Test Results

Variables	pov	rem	gdp	Δpov	Δrem	Δgdp
Null: Unit root						
LLC	-0.72810	-3.49010	0.49407	-8.04295	-9.10416	-6.95247
	(0.2333)	(0.0002)***	(0.6894)	(0.0000)***	(0.0000)***	(0.0000)***
Breitung t-stat	1.55492	2.47924	1.71590	-7.76180	-2.52354	-1.63297
	(0.9400)	(0.9934)	(0.9569)	(0.0000)***	(0.0000)***	(0.0000)***
IPS	1.37394	-1.69174	4.13771	-5.37744	-6.37202	-4.09694
	(0.9153)	(0.0453)**	(1.0000)	(0.0000)***	(0.0000)***	(0.0000)***
ADF-Fisher Chi-square	10.3652	28.0994	8.32752	55.1588	64.1623	50.6521
	(0.8469)	(0.0308)**	(0.9385)	(0.0000)***	(0.0000)***	(0.0000)***
PP-Fisher Chi-square	11.3512	48.1336	16.6697	96.7563	100.846	108.032
	(0.7873)	(0.0000)***	(0.4073)	(0.0000)***	(0.0000)***	(0.0000)***

Note: $^{*}p < 0.1$; $^{**}p < 0.05$; $^{***}p < 0.01$

Table 3: Panel cointegration tests results

Variables	Within dimension (panel statistics)		
Method	Test	Statistics	Prob.
pov, gdp, rem	Panel v-Statistic	-1.131834	0.8711
Pedroni(1999, 20014)	Panel ρ-Statistic)	1.162902	0.8776
	Panel PP-Statistic	-1.100568^{***}	0.0005
	Panel ADF-Statistic	-2.379847^{***}	0.0087
	Within dimension (individual statistics)		
	Test	Statistics	Prob.
	Group ρ-Statistic	2.116847	0.9829
	Group PP-Statistic	-1.544347^{*}	0.0613
	Group ADF-statistic	-2.586887^{***}	0.0048

Note: $^{*}p < 0.1$; $^{**}p < 0.05$; $^{***}p < 0.01$

Table 4: Granger causality tests results based on panel vector correction

Dependent Variables		Short-Term Causality		Long-Term Causality
	Δpov	Δrem	Δgdp	ECT(-1)
Δpov		6.682685**	7.830974***	−0.064551**
		(0.0354)	(0.0199)	(0.0262)
Δrem	2.795955		4.808658*	-0.018726
	(0.2471)		(0.0903)	(0.2768)
Δgdp	4.646179*	2.757747		-0.001407
	(0.0980)	(0.2519)		(0.2863)

Note: $^{*}p < 0.1$; $^{**}p < 0.05$; $^{***}p < 0.01$

Table 5: DOLS and FMOLS estimates of the long-run relationship

	DOLS	FMOLS
ln (gdp)	−0.78***	−0.72***
	0.00	0.00
ln (rem)	−0.15***	−0.13**
	0.01	0.00

Note: $^{*}p < 0.1$; $^{**}p < 0.05$; $^{***}p < 0.01$

Pedroni statistics

Panel V-statistic:

$$T^2 N^{3/2} Z_{\widehat{v}N,T} = T^2 N^{3/2} \left(\sum_{i=1}^{N} \sum_{t=1}^{T} \widehat{L}_{11i}^{-2} \widehat{\epsilon}_{i,t-1}^{2} \right)^{-1}$$

Panel ρ-statistic:

$$T\sqrt{N} Z_{\widehat{\rho}N,T^{-1}} = T\sqrt{N} \left(\sum_{i=1}^{N} \sum_{t=1}^{T} \widehat{L}_{11i}^{-2} \widehat{\epsilon}_{i,t-1}^{2} \right)^{-1} \sum_{i=1}^{N} \sum_{t=1}^{T} \widehat{L}_{11i}^{-2} (\widehat{\epsilon}_{i,t-1} \Delta \widehat{\epsilon}_{i,t} - \widehat{\lambda}_i)$$

Panel t-statistic (non-parametric):

$$Z_{tN,T} = \left(\tilde{\sigma}_{N,T}^{2} \sum_{i=1}^{N} \sum_{t=1}^{T} \widehat{L}_{11i}^{-2} \widehat{\epsilon}_{i,t-1}^{2} \right)^{-1/2} \sum_{i=1}^{N} \sum_{t=1}^{T} \widehat{L}_{11i}^{-2} (\widehat{\epsilon}_{i,t-1} \Delta \widehat{\epsilon}_{i,t} - \widehat{\lambda}_i)$$

Panel t-statistic (parametric):

$$Z_{tN,T}^{\star} = \left(\tilde{s}_{N,T}^{\star 2} \sum_{i=1}^{N} \sum_{t=1}^{T} \widehat{L}_{11i}^{-2} \widehat{\epsilon}_{i,t-1}^{\star 2} \right)^{-1/2} \sum_{i=1}^{N} \sum_{t=1}^{T} \widehat{L}_{11i}^{-2} \widehat{\epsilon}_{i,t-1}^{\star} \Delta \widehat{\epsilon}_{i,t}^{\star}$$

Group ρ-Statistic:

$$TN^{-1/2} \tilde{Z}_{\widehat{\rho}N,T^{-1}} = TN^{-1/2} \sum_{i=1}^{N} \left(\sum_{t=1}^{T} \widehat{\epsilon}_{i,t-1}^{2} \right)^{-1} \sum_{t=1}^{T} (\widehat{\epsilon}_{i,t-1} \Delta \widehat{\epsilon}_{i,t} - \widehat{\lambda}_i)$$

Group t-Statistic (non-parametric):

$$N^{-1/2} \tilde{Z}_{tN,T} = N^{-1/2} \sum_{i=1}^{N} \left(\widehat{\sigma}_i^{2} \sum_{t=1}^{T} \widehat{\epsilon}_{i,t-1}^{2} \right)^{-1/2} \sum_{t=1}^{T} (\widehat{\epsilon}_{i,t-1} \Delta \widehat{\epsilon}_{i,t} - \widehat{\lambda}_i)$$

Group t-Statistic (parametric):

$$N^{-1/2}\tilde{Z}^{\star}_{tN,T} = N^{-1/2}\sum_{i=1}^{N}\left(\sum_{t=1}^{T}\widehat{s}_i^{\star 2}\widehat{\epsilon}_{i,t-1}^{\star 2}\right)^{-1/2}\sum_{t=1}^{T}\widehat{\epsilon}^{\star}_{i,t-1}\Delta\widehat{\epsilon}^{\star}_{i,t}$$

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