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Ali, Abdilahi; Alpaslan, Baris

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IS THERE AN INVESTMENT MOTIVE BEHIND REMITTANCES? EVIDENCE FROM PANEL COINTEGRATION

Abdilahi Ali*
Aberystwyth University, UK
Baris Alpaslan
Yildrim Beyazit University, Turkey

ABSTRACT

Remittance flows have become a vital source of foreign exchange for many developing countries. As a result, the issue of whether they act as complements or substitutes for domestic investment remains an important avenue of research. We know that remittances can act as compensatory transfers, in which case altruistic motives may dominate. We also know that they can act as standard capital flows, where self-interest/ investment motivates may dominate. Hence, the motives behind remittance flows can have a direct bearing on how they influence domestic capital formation. In addition, the short-run relationship between domestic investment and remittances may be different from their long-run relationship. In light of these considerations, this paper re-examines whether migrant remittances "crowd in" or "crowd out" investment in developing countries, using a sample of 47 developing and emerging economies. The paper employs recently developed panel cointegration techniques given that these can overcome a number of important issues. First, we explicitly account for cross-sectional dependence, outliers as well as cross-sectional heterogeneity. Second, since our variables of interest may be influenced by various factors emanating from, for example, domestic policy changes or global economic trends, we account for structural breaks and regime shifts. Third, the approaches we employ are robust to endogeneity and many forms of omitted variable bias. Fourth, we examine both the long-run as well as the short-run relationship between remittance flows and domestic investment, employing panel error correction model to uncover the short-run dynamics. Finally, we conduct a panel Granger causality analysis to establish whether these relationships are indeed of a causal nature. The results of the paper show that remittances form a long-run equilibrium relation with domestic investment. The results of the panel vector error correction model reveal the absence of a short-run relationship but the presence of a long-run bidirectional link between remittances and investment. Thus, remittances drive investment while investment itself causes more remittances, suggesting that remittances are not only driven by altruistic motives but also investment motives. This long-run (causal) two-way relationship is robust to a battery of sensitivity analyses. However, when the sample is disaggregated into regions, the results of the Asian sub-sample are statistically insignificant. We suspect that this is due to the low number of observations from that region. An important policy implication emanating from this study is that developing countries should improve the effectiveness of remittance inflows given that these can augment the rate of capital accumulation.

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Corresponding Author's Email Address: abdi.ali@aber.ac.uk

INTRODUCTION

Over the last few decades, migrant remittances have taken a more prominent role in developing countries. As a result, the question of whether they crowd-in or crowd-out domestic investment has become an important policy issue. In general terms, the macroeconomic effects of remittances largely depend on whether they act as pure compensatory transfers or capital flows (Chami, Fullenkamp and Jahjah, 2005). In the first case, altruistic motives dominate in the sense that the migrant is concerned with the well-being of his/her relatives. In the latter case, though, self-interest dominates, such that the migrant retains some sort of ownership over the assets and thus uses remittances to finance investment projects. From the point of view of economic development, it is important to examine the degree to which this 'investment motive' exists in developing countries given the pivotal role capital accumulation plays in the process of economic development.

However, irrespective of the motive behind remittances, the response of the overall economy to increases in remittances could be either negative or positive. On the one hand, remittance flows can have negative effects on the recipient economy through their adverse influences on income distributions (Orrenius et al. 2010), household's labor supply and savings rates (Chami, Fullenkamp and Jahjah, 2005). In addition, similar to any other resource inflow, sustained levels of remittances tend to be associated with "Dutch disease" effects (Amuedo-Dorantes, Bansak and Pozo, 2005; Rahman, Foshee and Mustafa, 2013), output shocks (Imai et al. 2014) as well as increases in conspicuous consumption rather than productive investments (Chami, Fullenkamp and Jahjah, 2005).

On the other hand, there is considerable evidence showing that, although remittances may mainly go to consumption, a substantial portion of it goes to human capital formation in the form of better nutrition, schooling and health (Gupta, Pattillo, and Wagh, 2009). Moreover, increased consumption and even "unproductive" investments (e.g. real estate) can have significant multiplier effects, encouraging more capital accumulation and growth through spillover effects (Ratha, 2003; Gupta, Pattillo, and Wagh, 2009).

Evidence also suggests that remittances tend to reduce households' credit constraints and thus boost the depth of the financial sector (Guilamo and Ruiz-Arranz, 2009; Aggarwal, Demirguc-Kunt and Pera, 2011). Furthermore, it has been shown that remittance receiving households, on average, tend to save and invest more than other comparable households (Adams, 2006). Other studies found that remittances are associated with poverty reduction (Adams and Page, 2005) and higher educational attainments (Rapoport and Docquier, 2006). Finally, remittance flows have been found to act more counter-cyclically than other types of inflows and thus are a more stable source of foreign exchange at times of economic difficulties (Combes and Ebeke, 2011; Chami, Hakura and Montiel, 2009).

The objective of this study is to contribute to this literature but we depart from the existing literature in a number of ways. First, we use recently developed panel cointegration tests that can handle a number of econometric issues, including cross-sectional heterogeneity, structural breaks and endogeneity concerns. Second, we examine the long-run relationship between remittance inflows and domestic investment *directly* unlike studies such as Alleyne, Kirton and Figueroa (2008) who consider the relationship between remittances and per capita incomes. Third, we apply panel error correction methods to uncover the short-run dynamics in the relationship. Finally, we conduct a panel Granger causality analysis in order to establish whether the long and short-run effects are indeed of a *causal* nature.

The paper is organized as follows. Section 2 sets out the econometric analysis, presenting the techniques used as well as the findings while Section 3 concludes.

EMPIRICAL ANALYSIS

Basic Model and Data

To examine the relationship between remittance flows and domestic investment, we use a balanced panel[1]. Thus, the study is restricted to countries with consistent data so we end up with 47 developing and emerging economies over the period 1980-2006. The summary statistics as well as a description of the sample is summarized in Table A1 in the appendix.

Given that our overriding objective is to examine whether migrant remittances "crowd in" or "crowd out" domestic investment in developing countries, it is imperative to employ appropriate methods that would enable us to uncover the long-run as well as short-run dynamics in the relationship. In addition, we are interested in empirically exploring the presence or absence of either/ or both altruistic and self-interest/investment motives, as explained previously. To achieve these objectives, we employ panel cointegration, panel vector error correction and panel Granger causality methods. The basic empirical model takes the following form:

$$INV_{it} = \alpha_i + \gamma_{it} + \beta REM_{it} + \varepsilon_{it} \tag{1}$$

where α_i and γ_{it} are, respectively, country specific fixed and time effects, capturing any country-specific unobservable that are relatively stable over time and ε_{it} is the error term. INV_{it} is the share of investment in GDP for countries $i = 1, \ldots N$ and time periods $t = 1, \ldots T$, and REM_{it} is the share of remittances in GDP, both sourced from World Development Indicators (2011).

As is the standard norm in panel cointegration studies (see for example, Crowder and de Jong, 2011; Herzer and Grimm, 2012), equation (1) is a parsimonious specification that solely focuses on the bivariate long-run link between *REM* and *INV*. The validity of this specification, however, requires that the variables in (1) are nonstationary or, more precisely, integrated of the same order. In that case, they would have a stationary error term, implying that they constitute a cointegrating vector (Asteriou and Hall, 2007). Once a set of variables form a cointegrating relation, such (long-run) relationship should exist even if more variables are added to the model (see for example, Herzer and Grimm, 2012).

Panel Stationarity Tests

In estimating equation (1), we first test the time series properties of the variables using the panel unit root tests developed by Levin, Lin and Chu (2002) (*LLC*) and Im, Pesaran and Shin (2003) (*IPS*). The *LLC* is an extension of the standard (Augmented) Dickey-Fuller test and assumes parameter homogeneity while the *IPS* allows for heterogeneity across the panel and serial correlation in the error terms. Both the *LLC* and *IPS* may lead to erroneous results if there is cross-sectional dependence among the panel members emanating from, for example, common effects. Hence, we also report the cross-sectionally augmented panel unit root test (*CIPS*) proposed by Pesaran (2007), which takes into account possible cross-sectional dependence.

Table 1 reports the results of the unit root tests which indicate that we cannot reject the null hypothesis of a unit root in levels, suggesting that the variables are non-stationary. However, the series are stationary in first-differences, implying that they are integrated of order one, *I* (1). Hence, we can now proceed with panel cointegration tests to explore whether there is a long-run equilibrium relationship between *REM* and *INV*.

TABLE 1. PANEL UNIT ROOT TEST RESULTS

	LLC statistics		**IPS statistics**		**CIPS statistics**	
	Levels	**Diff**	**Levels**	**Diff**	**Levels**	**Diff**
Full sample						
$Investment_{it}$	-0.41	-1.22**	-2.21	-3.09**	-2.22	-2.73**
$Remittance_{it}$	-0.23	-1.05**	-1.41	-2.81**	-2.16	-2.70**
Africa						
$Investment_{it}$	-0.39	-1.25**	-2.09	-3.06**	-2.06	-2.54**
$Remittance_{it}$	-0.31	-1.15**	-1.90	-3.30**	-2.23	-2.78**
Latin America						
$Investment_{it}$	-0.51	-1.29**	-2.31	-3.23**	-2.41	-2.84**
$Remittance_{it}$	-0.06	-0.67	-0.72	-2.00**	-2.04	-2.20**
Asia						
$Investment_{it}$	-0.39	-1.07*	-2.37	-2.91**	-2.79**	-2.86**
$Remittance_{it}$	-0.28	-1.13	-1.52	-2.75**	-1.57	-2.68**

*Notes: The tests are: Levin, Lin and Chu (2002, LLC), Im, Pesaran and Shin (2003, IPS) and Pesaran (2007, CIPS . ** indicates the rejection of the null of non-stationarity at the 5% level or better. Two lags used to account for autocorrelation and the tests include intercept and trend in levels.*

Panel Cointegration Tests

Having established that the variables under study are *I* (1), we now explore whether there is a long-run cointegration between *INV* and *REM*. To this end, we implement the residual based panel cointegration test developed by Kao (1999) which is an ADF-type test. The null hypothesis tested here is that there is no panel cointegration against the alternative of cointegration based on the assumption of homogenous cointegrating vectors. Since the assumption of homogeneity among the cross-sectional units may be too strong, we also report the Pedroni (1999, 2004) panel cointegration test which offers considerable flexibility as it allows for heterogeneity in the long-run cointegrating vectors. Pedroni (1999, 2004) constructs seven test statistics which capture both the within- and between-dimensions of the panel.

However, an important shortcoming with the above panel cointegration tests is that they impose a common factor restriction - that is, they assume that the long-run parameters for the level variables are equal to the short-run parameters of the variables in their first differences. As shown by Westerlund (2007), when this assumption does not hold, the above cointegration methods suffer from a significant loss of power. Therefore, in addition to the above methods, we also report more appropriate panel cointegration tests proposed by Westerlund (2007). Westerlund (2007) sidesteps the assumption of a common factor restriction by utilizing the structural (rather than residual) dynamics. The Westerlund test can handle serially correlated residuals, country-specific intercept and slope parameters along with trend terms. Westerlund (2007) develops four different statistics which can be used to establish the existence of a panel cointegration. Two of them are *panel* tests (denoted by P_τ and P_α), testing the alternative hypothesis that the panel is cointegrated as a whole ($H_1^p := \alpha_i = \alpha < 0$ for all i). The other two are *group-mean* statistics, (denoted by G_τ and G_α), which test the alternative that at least one element in the panel is cointegrated ($H_1^g := \alpha_i < 0$ for at least one i). Thus, the panel tests assume that α_i is homogenous for all i while the group-mean tests do not require this.

To formally examine whether the panel members are indeed independent, we apply the *CD* test proposed by Pesaran (2004). Pesaran (2004) shows that the *CD* test is robust to a single as well as multiple breaks in the slope parameters and/or in the residual variances of the individual regressions.

Given the length of the time period we cover and the heterogeneity of the countries under study, it is highly likely that our variables of interest may have been influenced by various shocks emanating from, for example, regime and policy changes. Thus, to fully understand the relationship between *INV* and *REM*, structural breaks and regime shifts need to be accounted for. In this study, as an additional robustness, we implement the panel cointegration test proposed by Westerlund and Edgerton (2008), which accounts for both structural breaks and cross-sectional dependence. Westerlund and Edgerton (2008) develop two different tests that allow for unknown structural breaks in both intercept and slope of the cointegrating model, heteroskedastic and serially correlated errors as well as time trends. The location of the structural breaks may be at different dates for the cross-sectional units.

In the top panel of Table 2 below, we report the results of the Kao (1999) test which strongly rejects the null hypothesis of no cointegration between *INV* and *REM*. The null of no cointegration is also rejected when we allow for heterogeneous cointegrating vectors using the Pedroni (1999, 2004) tests. The table also reports the results based on Westerlund (2007). To account for cross-sectional dependence, bootstrapped robust *p*-values are reported (based on 500 replications). The results indicate that the null hypothesis of no cointegrating relationship can be rejected irrespective of whether we treat $\alpha_i$as homogenous (tests P_τ and P_α) or not (tests G_τ and G_α). Thus, there is a strong evidence of a cointegrating relationship between *REM* and *INV*.

To formally establish the existence of a cross-sectional dependence, we apply the *CD* test which strongly rejects the null hypothesis of no cross-sectional dependence (see Table 2). Thus, a failure to take this into consideration may result in biased results.

Finally, we consider the effects of structural breaks and regime shifts on the long-run relationship between *REM* and *INV* using the test developed by Westerlund and Edgerton (2008). Table 2 reports the results for three cases (no-break, level-break and regime-shift). When possible structural breaks are ignored (the no-break case) or accounted for (the level-break case), the null hypothesis of no cointegration can be rejected. However, when we consider regime shifts we fail to reject the null of no cointegration.

TABLE 2. PANEL COINTEGRATION TEST RESULTS – FULL SAMPLE

	Residual-based tests		
			T-statistics
Kao (1999)		ADF	-2.983***
Pedroni (1999, 2004)		Panel ν-stat	-5.235***
		Panel ρ-stat	-2.808***
		Panel PP-stat	-6.736***
		PanelADF-stat	-8.647***
		Group ρ-stat	1.087
		Group PP-stat	-3.585***
		Group ADF-stat	-5.480***
	Panel cointegration with cross-sectional dependence		
Westerlund (2007)		G*t*	-2.314**
		G*a*	-7.765***
		Pt	-14.221***
		P*a*	6.588***
	Cross-sectional independence tests		
Pesaran (2004)		CD without a linear trend	12.660***
		CD with a linear trend	12.010***
	Panel cointegration tests with structural breaks and CD		
Westerlund and Edgerton (2008)	Model	Z(N)τ	Z(N)φ
	No break	-11.530***	-20.550***
	Level break	-8.350***	-17.840***
	Regime shift	3.70	0.062

*Notes: The null hypothesis of the Kao and Pedroni tests is that the variables are not cointegrated and the lag lengths are based on Schwartz Information Criterion with a maximum number of 3 lags. Under the null, the Pedroni tests are distributed as normal and their finite sample distribution are tabulated in Pedroni (2004). For the Westerlund (2007) test, the optimal lag/lead length is determined by Akaike Information Criterion with the maximum of lags set equal to 3 and the width of Bartlett-kernel is set to 3 (bootstrapped robust p-values reported). The Pesaran (2004) CD test takes cross-sectional independence as the null and its associated p-values are for a one-sided test based on normal distribution. The lag length selection of the Westerlund and Edgerton (2008) test is based on an automatic procedure and 3 breaks are used based on grid search at the minimum of the sum of squared residuals. The p-values are for a one-sided test based on the normal distribution. ** denotes* *significance level at the 5% or better.*

The results hold when we divide the sample into (regional) sub-samples (see Table 3). So to sum up, we find that there is a long-run relationship between *INV* and *REM*. This link is robust to heterogeneity in the long-run cointegrating vectors as well as to cross-sectional dependence and structural breaks. Hence, we should have more confidence in the relationship between *INV* and *REM*. However, the results suggest that this link is not robust to regime shifts. With this in mind, we now estimate the nature of this relationship.

TABLE 3. PANEL COINTEGRATION TESTS: SUB-SAMPLES

		Kao (1999) test
		ADF
Africa		1.97**
Asia		-3.42**
Latin America		-4.05**
		Pedroni (1999, 2004) test
	Panel-PP	Group-PP
Africa	-2.32**	-3.27**
Asia	-1.02	-0.63
Latin America	-4.22**	-1.94*
	Westerlund (2007) cointegration test	
Africa	G*t*	-2.036**
	G*a*	-6.884**
	P*t*	-8.317**
	Pa	-5.511**
Latin America	G*t*	-2.948**
	G*a*	-10.029**
	P*t*	-9.179**
	Pa	-8.197**
Asia	G*t*	-1.979
	G*a*	-6.358**
	P*t*	-8.250**
	Pa	-8.131**
		Pesaran (2004) CD tests
	Without Trend	With Trend
Africa	3.09**	3.05**
Asia	9.47**	3.41**
Latin America	3.47**	8.50**

***Notes**: For an explanation of the tests, see notes under Table 2.*

Long-run Estimation

Having confirmed the presence of a cointegration, we apply the within-dimension-based dynamic OLS (*WD-DOLS*) estimator developed by Kao and Chiang (2001) to uncover the effect of *REM* on *INV*. To implement the WD-DOLS estimator, we consider the following panel model:

$$INV_{it} = \lambda_i + \beta REM_{it} + \varepsilon_{it} \tag{2}$$

Because our data is non-stationary, the *WD-DOLS* estimator addresses issues of serial correlation and endogeneity concerns by augmenting equation (2) with leads and lags of the first differences of the right hand side (endogenous) variable as follows:

$$INV_{it} = \lambda_i + \beta REM_{it} +_{j=-q}^{q} \Psi_{ij} \Delta REM_{it+j} + \upsilon_{it} \tag{3}$$

where Ψ_{ij} are the leads and lags. The *WD-OLS* estimator is super-consistent, under cointegration, producing unbiased estimates of the long-run cointegrating relationship.

Nevertheless, a particular weakness with the *WD-DOLS* estimator is that it assumes that the slope coefficients are homogenous across the cross-sectional units. However, this pooling assumption, if not true, can result in a serious bias in both static and dynamic panels (Asteriou and Hall, 2007). Thus, as a robustness check, we also estimate our model (equation 2) using the between-dimension mean-group DOLS (*MG-DOLS*) estimator for heterogeneous cointegrated panels suggested by Pedroni (2001). This estimator allows the long-run slope coefficients to vary across countries by running separate regressions for each cross-section and then averaging them, $\hat{\beta} = N^{-1}\sum_{i=1}^{N}\hat{\beta}_{it}$. Thus, the estimates can be viewed as the mean value of the individual cointegrating vectors. As emphasized by Pesaran and Smith (1995), group-mean estimators generate more consistent estimates, in the presence of heterogeneous cointegrating vectors, than do within-dimension estimators. In addition, the *MG-DOLS* estimator has better small sample properties (Pedroni, 2001).

As highlighted previously, we need to consider the possible issue of cross-sectional dependency. For example, investment rates and remittance flows in our sample of countries may respond to (unobserved) common external shocks (e.g. global business cycles), meaning that they may become correlated across i. Ignoring this interdependence may result in erroneous estimates. A simple way to deal with this type of error dependence is to demean the data over the cross-sectional units so that the cross-section averages of the variables, say $\bar{x_t} = N^{-1}\sum_{i=1}^{N} x_{it}$ are subtracted from the observations, say x_{it}. This procedure can mitigate the effects of error dependence (Pedroni, 2001; Levin et al. 2002). Thus, we re-estimate the *WD-DOLS* regressions using demeaned data. This simple strategy, while effective, implies that the unobserved external factors are the *same* across countries. To the extent that countries have different macroeconomic and institutional environments, for example, it is highly likely that their responses and behavior towards remittances would be different. To this end, we also apply the Common Correlated Effects Mean Group estimator (*CCEMG*) developed by Pesaran (2006). Applying this estimator, one can rewrite the error term in Equation (2) as having a multifactor structure as follows:

$$\varepsilon_{it} = \acute{\omega}_i \Pi_t + \upsilon_{it} \tag{4}$$

where Π_t is *k x 1* vector of unobserved common factors, which may affect the countries with different intensities, υ_{it} is country-specific error term, assumed to be weakly dependent across the cross-sectional units. The common factors Π_t are allowed to be correlated with the regressors in Equation (2):

$$x_{it} = \eta_i + \acute{\xi}_t \Pi_t + \varepsilon_{it} \tag{5}$$

where x_{it} is each of our regressors, $\acute{\xi}_t$ is *k x 1* vector of loadings, and ε_{it} is the error term assumed to be independently distributed of Π_t and υ_{it} .

To take into account the presence of common effects, Pesaran (2006) suggests that one can approximate Π_t by cross-section averages of the dependent and explanatory variables and then run standard panel regressions augmented with these averages. As shown by a number of studies (e.g. Pesaran, 2006; Pesaran and Tosetti, 2011), this *CCEMG* performs well in small samples and can handle the presence of autocorrelation in the residuals and unit roots in the common factors.

As a final robustness check, we apply Breitung's (2005) two-step estimator which, unlike the above methods, can handle dynamic effects. Following Breitung (2005), it can be shown that a cointegrated model has the following Vector Error Correction Model (*VECM*) representation (in the case of a *VAR[1]*):

$$\Delta y_{it} = \alpha_i \beta' y_{it-1} + \varepsilon_{it} \tag{6}$$

where ε_{it} is a white noise error with $E(\varepsilon_{it}) = 0$ and positive definite covariance matrix $\sum_i = E(\varepsilon_{it}\varepsilon_{jt})$. The matrix β' captures the long-run relationship among the variables and is assumed to be the same across i while α_i and $\sum_t$ are short-run parameters which vary across i. In the first step, the country-specific short-run parameters are generated from separate models for each cross-section unit resulting in country-specific cointegration vectors. In the second step, the long-run cointegration matrix β' is estimated using the pooled regression:

$$\hat{q}_{it} = \beta' y_{it-1} + \hat{\upsilon}_{it} \tag{7}$$

where $\hat{q}_{it}$ and $\hat{\upsilon}_{it}$ are based on the generated short-run parameters α_i and $\sum_t$. Breitung (2005) and Breitung and Pesaran (2008) show that this estimator has a normal distribution and corrects for endogeneity in the second step.

Table 4 contains the results of the estimates of the long-run effects of *REM* on *INV*. The coefficient of *REM* is positive and highly significant at the 1% level. The magnitude of the coefficient ranges between 0.22 and 0.63, implying that, in the long-run, a one percentage point increase in the *REM* to GDP ratio leads to an increase in *INV* of around 0.22 – 0.63 percentage points. We observe similar results when we disaggregate the sample into 3 regional groups. In particular, the coefficient of the remittance variable is positive for all regions, albeit insignificant for the Asian countries. We suspect that the low number of Asian countries in the sample may be the cause of the insignificance of this variable.

TABLE 4. THE IMPACT OF REM ON INV

Estimator	$\mathbf{REM_{it}}$	*N*	*Obs*
WD-DOLS (Kao and Chiang, 2001)	0.431 [4.460]***	47	1269
WD-DOLS (Demeaned data)	0.222 [1.910]**	47	1269
MG-DOLS (Pedroni, 2001)	0.628 [9.380]***	47	1269
CCEMG estimator (Pesaran, 2006)	0.222 [0.981]	47	1269
2-step estimator (Breitung, 2005)	0.302 [6.293]***	47	1269
	Sub-samples#		
Africa	0.414 [3.24]***	21	567
Latin America	0.526 [2.66]**	15	405
Asia	0.151 [0.57]	11	297

Notes: *T-statistics* in parenthesis. ** and *** indicate significance at the 5% and 1% levels, respectively. The DOLS regressions are estimated with two leads and two lags. The regressions include unreported fixed effects. # Sub-sample estimates are based on the WD-DOLS estimator.

Short-run Dynamics and Causality Tests

Given that the variables are cointegrated, we set up a panel vector error correction model in order to explore whether the relationship between *REM* and *INV* is of a causal nature. To this end, following Engle and Granger (1987), we use the following two-step procedure (Pesaran, Shin and Smith, 1999). First, the long-run model specified in equation (2) is estimated in order to obtain its residuals. Second, defining the lagged residuals from equation (2) as the error correction term, the following error correction model is generated:

$$\Delta INV_{it} = \alpha_{ij} + \sum_{k=1}^{p} \gamma_{11ik}\,\Delta INV_{it-k} + \sum_{k=1}^{p} \gamma_{12ik}\,\Delta REM_{it-k} + \lambda_{1i}\varepsilon_{it-1} + \epsilon_{1it}, \tag{8}$$

$$\Delta REM_{it} = \alpha_{2j} + \sum_{k=1}^{p} \gamma_{21ik}\,\Delta REM_{it-k} + \sum_{k=1}^{p} \gamma_{22ik}\,\Delta INV_{it-k} + \lambda_{2i}\varepsilon_{it-1} + \epsilon_{2it}, \tag{9}$$

where Δ is the first-difference operator; *p* is the optimal lag length determined by standard information criterion. The null hypothesis of no short-run causality can be examined, respectively, based on $H_0\colon \gamma_{12ik} = 0$ and $H_0\colon \gamma_{22ik} = 0$ for all ik. In other words, short-run causality can be tested evaluating the statistical significance of the partial *F*-statistic associated with the corresponding regressor. On the other hand, long-run causality can be tested by the statistical significance of λ_{1i} and λ_{2i} (the error correction terms), respectively, using *T*-statistics.

The long and short-run Granger causality tests are reported in Table 5. The results show there is no *causal* relationship between *REM* and *INV* in the short-run for the sample as a whole or for the regional groups. However, in the long-run, we find a significant two-way causal relationship for the full sample as well as for the African and Latin American sub-samples. That is, increases in *INV* are both *a result of* as well as *a cause of* increases in *REM*. This suggests that remittances are not only driven by altruistic motives but *also* investment motives.

TABLE 5. SHORT-RUN DYNAMICS AND CAUSALITY TESTS

Dependent variable	**Source of causality**		
	Short-run		**Long-run**
			ECT
Equation (8)	-	1.260 [0.262]	0.458*** [0.000]
Equation (9)	1.920 [0.166]	-	0.019*** [0.009]
		Africa	
Equation (8)	-	0.020 [0.893]	0.749*** [0.000]
Equation (9)	1.590 [0.207]	-	-0.008** [0.017]
		Latin America	
Equation (8)	-	3.640* [0.262]	0.400*** [0.000]
Equation (9)	0.190 [0.665]	-	0.026** [0.032]
		Asia	
Equation (8)	-	0.040 [0.851]	-0.210 [0.440]
Equation (9)	0.840 [0.360]	-	0.005 [0.674]

Notes: Partial F-statistics are reported with respect to short-run changes in the respective regressor. The ECM is the coefficient of the error correction term. *** indicates significance at the 1% level.

Additional Robustness

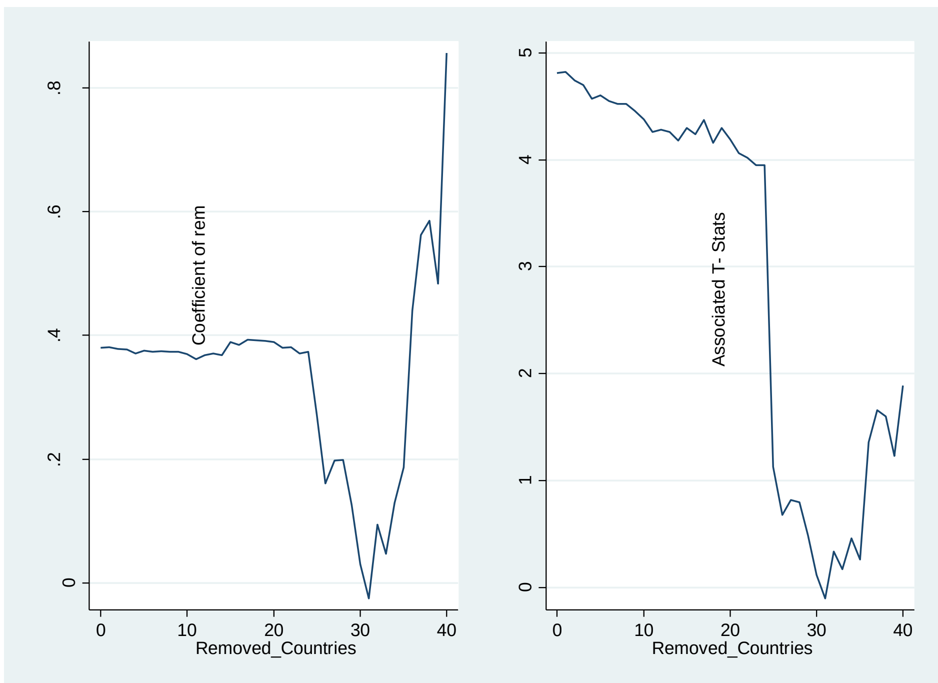
In the first instance, we explore whether the long-run results for the panel as a whole are sensitive to potential outliers. To this end, we re-estimate the main regression, removing one country at a time from the estimation. Figure 1 graphically shows the coefficients as countries are removed and their t-statistics. As can be seen, the coefficients of remittance always carry a positive sign which is mostly significant. Thus, the long-run results are robust to the exclusion of outliers.

As shown previously, our results are robust to endogeneity and omitted variable bias given the super-consistent properties of panel cointegration techniques. To formally show that this is indeed the case, we rewrite equation (1) in the form of a multivariate ARDL (*p,1,…q*) model. This would enable us to capture how domestic investment adjusts to changes to remittance flows and other regressors:

$$INV_{it} = \alpha_i + \sum_{k=1}^{p} \gamma_{ik}\, \mathrm{INV}_{it-k} + \sum_{k=0}^{q} \varsigma'_{ik}\, \mathrm{x}_{it-k} + \varepsilon_{it}, \tag{10}$$

where *i* and *t* index country and time, respectively, *k* is the lag length, α_i is a fixed effect and $\mathrm{x}_{it} = (REM_{it}, Controls_{it})$. We control for economic performance (real GDP per capita growth), macroeconomic stability (proxied by rate of inflation), openness (trade/GDP), and Foreign Direct Investment (FDI/GDP). According to Adams and Klobodu (2016), the impact of remittances on economic performance may be conditional on the presence of a democratic government. Following their study, we use Polity2 as a measure of democracy (proxy for regime type) and the variable ranges between –10 to 10, with higher values indicating a more democratic regime (Marshall, Gurr and Jaggers, 2014). As shown by Shimada (2012), the stock of international migration (as share of the population) can matter both for the level of remittances as well as their motivation. Hence, we also control for migration outflows[2]. All the variables are drawn from WDI (2011).

FIGURE 1. SENSITIVITY TO OUTLIERS



Note: This figure shows the coefficient of REM (left side) as countries are removed and their T-statistics (right side). These estimates are based on the DOLS estimator.

Following Pesaran and Smith (1995), equation (10) can be re-parameterized as:

$$\Delta INV_{it} = \alpha_i + \phi_i \mathrm{INV}_{it-1} + \beta'_i x_{it} + \sum_{k=1}^{p-1} \gamma_{ik}^{*} \Delta \mathrm{INV}_{it-k} + \sum_{k=0}^{q-1} \zeta_{ik}^{*\prime} \Delta \mathrm{x}_{it-k} + \varepsilon_{it}, \tag{11}$$

where,

$$\phi_i = -(1 - \sum_{k=1}^{p} \gamma_{ik}), \beta_i = \sum_{k=0}^{q} \zeta_{ik},$$

$$\gamma_{ik}^{*} = -\sum_{m=k+1}^{p} \gamma_{im}\,, k = 1,2, \dots p-1, and$$

$$\zeta_{ik}^{*} = -\sum_{m=k+1}^{q} \zeta_{im}\,, k = 1,2, \dots q-1,$$

To estimate the short-run dynamics as well as the long-run effects, equation (11) can be expressed as an error-correction model as follows:

$$\Delta INV_{it} = \alpha_i + \phi_i(\mathrm{INV}_{it-1} - \varphi'_i x_{it}) + \sum_{k=1}^{p-1} \gamma_{ik}^{*}\,\Delta\mathrm{INV}_{it-k} + \sum_{k=0}^{q-1} \zeta_{ik}^{*\prime}\,\Delta\mathrm{x}_{it-k} + \varepsilon_{it}, \tag{12}$$

where $\varphi_i = (\beta_i/\phi_i)$ captures the long-run relationship between investment and the regressors while γ_{ik}^{*} and $\zeta_{ik}^{*\prime}$ capture the short-run dynamics linking investment to both its past values and the other variables in the model. The error-correction parameter,ϕ_i, measures the speed of adjustment of investment to its long-run equilibrium following a change in the regressors. Provided that ϕ_i is significant and negative, one can deduce that the variables exhibit a return to long-run equilibrium (i.e. there is a long-run relationship between them).

The above ARDL specification overcomes issues of endogeneity since all the regressors enter the model with lags. In addition, it allows the parameters to be different for each country. To estimate equation (12), we use three alternative estimators that are suitable for nonstationary heterogeneous panels which were advanced by Pesaran and Smith (1995) and Pesaran, Shin and Smith (1999). The first is the Mean Group (MG) estimator which runs separate models for each i and then averages the coefficients, thus allowing separate intercepts, slope coefficients and error variances. The second is the Dynamic Fixed Effects (DFE) estimator which treats the slope coefficients and error variances to be equal across i, while allowing country-specific intercepts. The DFE also imposes the restriction that the speed of adjustment and short-run coefficients to be equal.

Finally, we have the Pooled Mean Group (PMG) estimator, which allows heterogeneous intercepts, error-correction terms and error variances but treats the long-run parameters to be the same across the countries. As suggested by Pesaran and Smith (1995), one can use Hausman test to check the validity of the long-run parameter homogeneity. Given that, in our case, the test fails to reject the null hypothesis of the homogeneity restriction, the PMG estimator produces more efficient and consistent estimates relative to the other two estimators. So the PMG is our preferred specification.

The results are reported in Table 5. The first thing to notice is that the error-correction term is negative and highly significant. This indicates that, in line with our previous findings, we have stationary residuals and hence a non-spurious long-run equilibrium relationship among the variables. Also, this justifies the parsimonious specification we adopted earlier and is in line with the well-known cointegration proposition, which states that, provided there is cointegration between (two) variables, such (long-run) relationship should exist even if more variables are added to the specification (see for example, Herzer and Grimm, 2012).

TABLE 6. LONG AND SHORT-RUN DETERMINANTS OF INVESTMENT

	Pooled Mean Group		Dynamic Fixed Effects	
Variables	**[1]**	**[2]**	**[1]**	**[2]**
		Long-Run Estimates		
Remittance	0.277**	0.230*	0.288**	0.289**
	(0.114)	(0.118)	(0.129)	(0.130)
FDI	0.182	0.150	0.118	0.114
	(0.130)	(0.135)	(0.210)	(0.211)
Trade openness	0.077***	0.067***	0.058*	0.057*
	(0.020)	(0.020)	(0.030)	(0.031)
Inflation (logs)	-0.046	-0.105	0.260	0.267
	(0.200)	(0.208)	(0.463)	(0.466)
GDP Growth	1.066***	1.130***	0.850***	0.859***
	(0.087)	(0.095)	(0.277)	(0.279)
Polity2	-0.229***	-0.213***	-0.110	-0.104
	(0.051)	(0.049)	(0.086)	(0.086)
Migrant stock (logs)		-1.545		-1.242
		(3.941)		(7.837)
Error Correction Coefficients				
Phi	-0.299***	-0.292***	-0.301***	-0.300***
	(0.026)	(0.026)	(0.075)	(0.075)
		Short-Run Estimates		
D(Remittance)	-3.765	-2.570	-0.105	-0.106
	(3.574)	(2.320)	(0.099)	(0.098)
D(FDI)	0.200	0.418**	0.253***	0.254***
	(0.160)	(0.189)	(0.074)	(0.074)
D(Trade openness)	0.030	0.038	0.046	0.046
	(0.028)	(0.027)	(0.028)	(0.029)
D(Inflation, logs)	0.140	0.199	-0.010	-0.013
	(0.158)	(0.187)	(0.124)	(0.124)
D(GDP Growth)	-0.051	-0.057	-0.073*	-0.076*
	(0.039)	(0.040)	(0.042)	(0.042)
D(Polity2)	0.056	-0.134	0.057***	0.056
	(0.317)	(0.454)	(1.224)	(0.042)
D(Migrant stock, logs)		43.824		2.697
		(29.888)		(2.312)
Intercept	3.379	4.690	3.740	4.677
	(0.427)	(0.543)	(1.224)	(6.288)
Observations	1041	1041	1041	1041

*Notes: Standard errors in brackets. The estimates based on the MG estimator are not reported given that the calculated Hausman statistic is, respectively, 1.98 and 1.39 for models [1] and [2]. The test statistic is distributed as chi2(6) and chi2(7), respectively. The regressions control for country and time effects * **, ** and * denote significance level at the 1, 5 and 10% respectively.*

Focusing on our variables of interest, the long-run estimates confirm the robustness of our previous results. That is, remittances are positively and significantly linked to domestic investment. The other variables are broadly in line with our a priori expectations, except the Polity variable which produces mixed results. The migrant stock variable is also insignificant. Consistent with our previous panel Granger causality results, remittances do not have any short-run relationship with investment.

Discussion of the Findings

Our central findings show that remittances have a robust long-run effect on domestic investment in developing countries. This result is consistent with the recent findings by Ziesemer (2010), who has shown that remittances enhance fixed capital formation directly as well as indirectly through their beneficial influences on public expenditures on education and literacy. The idea that remittance flows improve human capital (e.g. education, nutrition and health) has been confirmed by a number of studies (see for example, Acosta et al. 2007; Calero et al. 2009). Hence, these flows are likely to have positive effects in the long-run. Our findings are also in line with the results of Nsiah and Fayissa (2013) who found that remittances are positively related to economic development in developing countries. Similar to us, these authors found that the positive impact is significant in Latin America and Africa, while this is not the case in Asian countries. Unlike their study, however, we pay particular attention to the properties of the variables under study as well as the underlying assumptions of the econometric techniques. Given that we employ more superior estimation methods, our results should be more reliable.

As shown throughout, our results are not sensitive to cross-sectional heterogeneity, structural breaks, outliers and endogeneity issues. It should be emphasized that long-run positive effects of remittances on investment remain even when we account for democratic regime and migration outflows.

Our causality analysis shows that there is a bidirectional causal relationship between *REM* and *INV*. In particular, remittances drive investment while investment itself causes more remittances, suggesting that remittances are not only driven by altruistic motives but also investment motives. This could be because of the multiplier effects generated by the expenditures of remittance-receiving households may be encouraging more investment. Alternatively, it could be that the households themselves may be making small capital investments. In the latter case, this could generate more remittance flows if we assume that the migrant is not just altruistic but also self-interested. In other words, if remittance-receiving households engage in successful business ventures, the migrants may send more remittances in order to enhance their own wealth provided adverse selection and moral hazard can be overcome and trust can be established. Results by Alleyne et al. (2008) confirm that remittances are not only driven by altruistic motives but also investment motives. Thus, remittances may drive investment while investment itself may cause more remittances. These ideas are consistent with the theoretical work by Le (2011), who has shown that remittances can act as a useful source of finance for investment projects particularly when the domestic financial system is sufficiently developed.

Throughout, we emphasized two main rationales for remittance flows based on the seminal contributions by Chami et al. (2005) – namely, investment and altruism motives. However, Amuedo-Dorantes and Pozo (2006) highlight a third motive – namely, risk sharing and coinsurance. Using Mexico as a case study, they show that migrants may remit more when they face income and other risks in the host economy. Our results are consistent with this risk-coverage and insurance motive as well. Hence, remittances are not only driven by altruistic motives but also investment as well as insurance motives[3].

CONCLUDING REMARKS

The objective of this study was to establish whether there is a long-run stable relationship between domestic investment and remittances in developing countries. Using recently developed panel cointegration techniques that take into account issues such as cross-sectional dependence, heterogeneity, omitted variable bias as well as endogeneity and structural and regime changes, we show that there is a long-run relationship between investment and remittances. In particular, remittance flows are positively related to investment and thus economic development. This is largely in line with the findings of, among others, Imai et al. (2014), Ramirez and Sharma (2008) and Adams (2006).

The results of the panel vector error correction model reveal the absence of a short-run relationship but the presence of a long-run bidirectional link between remittances and investment. Thus, remittances drive investment while investment itself causes more remittances, suggesting that remittances are not only driven by altruistic motives but also investment motives. This long-run (causal) two-way relationship is robust to a battery of sensitivity analyses.

The overall findings suggest a number of important policy implications. First, developing countries should improve the effectiveness of remittance inflows. A particular channel is the financial system. Thus, developing countries should develop their financial sectors in order to allow remittance-receiving households to have the facilities needed for productive investments. Given that remittances tend to boost the level of deposits and credit in banking system (Aggarwal et al. 2011), a well-developed financial system would likely generate more benefits. In the same vein, they should adopt policies that may reduce the transaction costs attached to receiving the funds so that households can get their remittances as smoothly as possible. One way to do this is to reduce red tape, but perhaps, more importantly, competition should be encouraged among money transfer companies.

Overall, the important role migrant remittances can play in economic development is not a trivial matter. As shown in this study, remittances can improve the economic performance of developing economies by augmenting the rate of capital accumulation.

APPENDIX

TABLE A1. SUMMARY STATISTICS AND SAMPLE DESCRIPTION

	Mean	Standard error	Observations
Investment	21.60	7.96	1269
Remittances	4.20	9.44	1269
FDI	1.85	4.13	1269
Polity2	0.96	6.66	1215
Trade openness	67.96	35.90	1267
Inflation (logs)	3.03	0.75	1265
Migration stock (logs)	0.63	1.19	1269
Real per capita GDP growth	3.47	4.72	1268

Notes: Countries in the sample: Algeria, Argentina, Bangladesh, Bolivia, Botswana, Brazil, Burkina Faso, Cameroon, Colombia, Costa Rica, Cote d'Ivoire, Dominica, Dominican Republic, Egypt, El Salvador, Gabon, The Gambia, Ghana, Guatemala, Honduras, India, Jordan, Kenya, Lesotho, Madagascar, Mauritania, Mexico, Morocco, Mozambique, Niger, Pakistan, Panama, Papua New Guinea, Paraguay, Philippines, Rwanda, Senegal, Sri Lanka, Sudan, Suriname, Syria, Thailand, Togo, Trinidad and Tobago, Tunisia and Turkey (N = 47). We refer to this sample as 'developing and emerging economies' based on IMF classification (see for example; Nielsen, L. (2011). Classifications of countries based on their level of development: How it is done and how it could be done. *IMF Working Papers*, 1-45 (page 19).

ENDNOTES

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[1] The techniques we are employing require a balanced panel so our sample comprises of developing countries with consistent data. Arguably, this may give rise to selectivity bias but given that the sample we end up with is randomly selected, any potential selectivity bias would likely be minimal.

[2] We owe the suggestions to control for both regime type and migration outflows to the referee.

[3] To examine the underlying motives more methodically, a distinction should be made between a theory of altruism and a theory of insurance but this is beyond the scope of this paper. We are grateful to the referee for highlighting this issue.

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