Remittances and Domestic Output in the Long Run: Theory and New Evidence*

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Abstract

We employ a heterogeneous cointegration model to investigate the long-run relationship between remittances and real output across 80 developing countries. The empirical model is motivated by a bivariate long run relation derived from a two-sector open-economy model with tradable and non-tradable goods where remittance is an important source of income to households. The coefficient of remittance in the long-run relation, which we allow to vary across countries captures a trade-off impact of remittance on steady state investment, consumption, labor supply, and trade balance. The results uncover that, on average, a 1 percent increase in remittances is associated with approximately 0.07 percent increase in output in the long-run. Nonetheless, this positive average relationship masks a large heterogeneity across countries. More precisely, the associated response of real GDP following a permanent increase in remittance varies from 0.59 percent in Dominican Republic to -0.53 percent in Bosnia and Herzegovina. We find that cross-country differences in financial development and institutional quality explain the observed heterogeneity in the remittance-output relationship.

JEL Classification: E1, E20, F24, F43
Keywords: Remittances, Output, heterogeneous panel cointegration, open-economy, Developing economies

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*We have NO conflicts of interests. We have received helpful comments and suggestions from Stephen O'Connell, David McKenzie, Markus Eberhardt, Justin Kakeu, John Hoddinott, Juan Munoz, Alberto Posso, Kwabena Gyimah-Brempong, Anne Barthel, and Eric Hoffmann that have significantly improved the paper. We also thank conference participants and discussants at the 2019 North East Universities Development Consortium (NEUDC) meeting at Northwestern University for useful discussions and feedback. We are solely responsible for any remaining errors.

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1 Introduction

"Remittances are dollars wrapped with care" — Dilip Ratha

According to the 2019 World Bank's Migration and Development Brief, annual remittances to low and middle-income countries reached a record high of \$529 billion in 2018—a 9.6 percent jump from the 2017 record. In the last decade, remittances have exceeded foreign aid to developing countries and are catching up to foreign direct investment (Barajas et al., 2018; Clemens and McKenzie, 2018).[1] Although aid and foreign direct investment remain important sources of capital flow in developing countries, particularly in fragile economies, the household-to-household nature of remittances makes them a unique source of revenue (Chami et al., 2008). Indeed, the United Nations' Sustainable Development Goals (SDGs) identify remittances as a lifeline for many families and communities as well as a potential driver of growth and development in developing countries.[2]

Interestingly, the links between remittance and development at the aggregate level are, both theoretically and empirically, ambivalent (Clemens and McKenzie, 2018; Faini, 2007). From a theoretical perspective, the economy-wide effect of remittances on output is transmitted through several channels including but not limited to labor supply, investment, human capital, consumption, and total factor productivity. Consequently, the aggregate effect of remittances on output depends critically on the sign and size of the impact of remittances on these mediating factors, which by themselves, have a complex relationship with real output at the cross-country level (Hulten and Isaksson, 2007). On the empirical front, econometric issues ranging from measurement errors, cross-sectional dependence due to common shocks, and endogeneity issues arising from omitted variables make obtaining reliable estimates for the relationship between remittances and growth challenging. Moreover, existing empirical studies that have attempted to address these econometric challenges when estimating the remittance-output relationship have done so under strong homogeneity assumptions— i.e., a common remittance-output relationship across countries. This paper explores the long-run remittance-output relationship under a more relaxed assumption of a heterogeneous remittance-output relationship across countries while attempting to address the aforementioned econometric challenges.

We combine theory and empirics to uncover the long-run relationship between remittances and output in developing countries. For the theory, we embed remittance in a two-sector real business open-economy model that incorporates a number of channels through with remittances impact output (with tradable and non-tradable goods), and derive a closed-form long-run relation between remittances and real GDP in the economy. In the long run relation, the

[1]It is worth mentioning that excluding China, remittances to low- and middle-income countries (\$462 billion) were significantly larger than foreign direct investment flows in 2018 (\$344 billion). See, https://www.worldbank.org/en/news/press-release/2019/04/08/record-high-remittances-sent-globally-in-2018 for more details on the World Bank Migration and Development Brief.

[2]See, https://www.un.org/en/events/family-remittances-day/un-action.shtml a discussion on the importance of remittances in achieving the SDGs at both the micro- and macro-level.

coefficient of remittance is the total effect of remittance on real GDP, which captures a trade-off impact of remittance on steady state level of investment, consumption, labor supply, and trade balance. Specifically, in the model, a permanent increase in remittances raises the long-run wealth of the country, and with constant prices in steady state, there is a permanent increase in traded and non-traded consumption generating a rise in aggregate consumption. However, the increase in consumption of non-traded goods implies an increase of output in non-traded goods. This leads to a reallocation of labor from the traded sector to the non-traded sector thereby inducing a fall in output in the traded sector. Moreover, the long run impact of remittances on investment is ambiguous and critically depends on whether labor is moving from a relatively less capital intensive sector to a relatively more capital intensive sector. In summary, the theory reveals that depending on model parameters, the net (total) long run effect of remittances on output can vary in size and even in sign across countries suggesting a potential heterogeneous long-run output-remittance relationship across countries.

Empirically, we employ data from 80 developing countries over the period 1974-2014 and utilize recently developed heterogeneous cointegration-panel techniques to a estimate a bivariate long run relationship between remittance and real GDP. An advantage of our bivariate cointegration specification is that the coefficient of remittance in the regression model captures the total effect of remittance on output as motivated by the theory. Intuitively, the estimated coefficient of remittance captures the total impact of remittances on output via its countervailing effects on investment, consumption, and labor supply. The bivariate setting also implies that we do not preclude any potential channel through which remittance can affect output. To further capture a key tenet of the theory, the empirical work allows for a heterogeneous remittance-output relationship across countries. More importantly, the cointegration estimation techniques we employ account for different forms of endogeneity arising from omitted variables and measurement error in the data— issues that generally make it difficult to detect a positive effect of remittance on output. Furthermore, we explicitly deal with cross sectional dependence that may arise from global shocks such as oil shocks, global financial crisis, and the ongoing COVID-19, as well as, economic spillovers in the empirical estimation. The advantages offered by our bivariate setting are not shared with the methodologies used in the existing studies (e.g. Chami et al., 2008; Fayissa and Nsiah, 2010; Giuliano and Ruiz-Arranz, 2009; Pradhan et al., 2008; Rao and Hassan, 2011).

We find that there exists a positive and statistically significant long-run relationship between remittances and output. In particular, our baseline estimate reveals that, on average, a one percent increase in remittance is associated with a 0.066 percent increase in real GDP in the long run. This result is robust to alternative estimators, potential outliers, sampling bias, model specification, and periodisation. Specifically, the robustness exercises uncover a consistently positive and statistically significant long-run relationship between remittances and output. Beyond the panel results, we find a strong heterogeneity in the size and the sign of the estimated output-remittance relationship across countries. Specifically, the associated response of real GDP following a permanent increase in remittance varies from -0.53 percent in Bosnia and Herzegovina to 0.59 percent in Dominican Republic. We find that this observed variation in the output-remittance relationship across countries can be explained by the level of institutional quality and financial development across countries. More precisely, we observe an inverted U-shaped relationship between overall financial development and the estimated remittance-output. The inverted U-shape relationship suggests that at early stages, financial development may be positively associated with the impact of remittances on output and vice-versa. In contrast, there exists an increasing relationship between institutional quality and the estimated impact of remittance on output. In particular, better institutions proxied by the ability of the government to formulate and implement sound policies and regulations that permit and promote private sector development is related to a stronger positive impact of remittance on output.

Finally, we pursue a narrower Granger causality approach for our panel, which allows for heterogeneous causal relationship across countries. The results reveal that there exists a long-run bi-directional causal relationship between remittance and output. This implies that a rise in remittance increases output and that, in turn, higher output induces an increase in remittance inflows. An important and direct implication of these findings is that the ongoing policy efforts to lower transaction costs to ease the transfer of remittances (see for example, World Bank, 2017), can generate a permanent and positive impact of remittance on real GDP. This can contribute non-trivially to sustainable development in these developing economies.

The rest of the paper is organized as follows: Section 2 places the contribution of the study in the literature. Section 3 introduces the model to lend a structural interpretation to our econometric specification. Section 4 presents the econometric methodology, Section 5 presents and discusses the baseline results. Section 6 conducts robustness exercises. Section 7 presents the plausible explanations for the heterogeneous relationship between remittances and output across countries. Section 8 uncovers the causal relationship between remittances and output. Section 9 concludes.

2 Related Literature

In the quest to understand the aggregate impact of remittance, the existing literature has predominantly focused on growth analysis. The results have generally been mixed, with findings ranging from small positive, zero, and negative effects (e.g. Chami et al., 2008; Giuliano and Ruiz-Arranz, 2009; Pradhan et al., 2008; Rao and Hassan, 2011). Existing studies have attempted to explain these mixed results by arguing that a positive remittance-growth effect is likely to be offset by, among other things, the negative impact of remittance on labor supply (e.g., Posso, 2012) or the components of multi-factor productivity such as institutions (Abdih et al., 2012).[3] More recently, Clemens and McKenzie (2018) succinctly explain three new reasons for why a positive remittance-growth relationship may be difficult to detect. The reasons include: (1) Endogeneity bias due to the omitted variable, rising emigration. This variable is correlated with rising remittances, and it is also an opportunity cost to GDP, (2) the possible lack of power of the cross-country growth regressions to detect growth effects of remittances, and (3) Measurement errors in remittance primarily driven by changes in the measurement of remittance. The authors show that these three issues can individually expunge the positive effect of remittance on output in empirical studies independent of prior explanations. Thus, despite providing great insight into the remittance-output nexus, existing empirical studies are likely to be subjected to at least one of these three new issues. We take the issues laid out in Clemens and McKenzie (2018) seriously and attempt to account for them in our study.

The cointegration techniques we employ allows us to use level data, which in turn permits us to utilize the rich and informative non-stationarity properties of output and remittance. This empirical strategy naturally removes the drawback of using growth rates. Studies that use cross-country growth regressions to quantify the impact of remittance on growth may lack power to detect a positive effect due to the high variance feature in measured growth and remittances (Clemens and McKenzie, 2018). Growth is a complex process that depends on a host of variables other than the inflow of external funds (Herzer and Morrissey, 2013). Consequently, to study the impact of remittance on growth, one needs to control for a host of potential determinants of growth such as investment and institutional quality, which may not only lead to overcontrolling but also induce the risk of precluding other potential channels through which remittance can affect growth. Furthermore, growth rates show little persistence over time while remittances, in levels or relative to GDP, have exhibited persistent trends.[4] A regression involving growth (which is stationary) as the dependent variable and the remittances/GDP ratio (which is non-stationary) as an independent variable is therefore unbalanced and can lead to misleading results (see Ericsson et al., 2001; Herzer and Morrissey, 2013).[5] This unbalanced analysis is avoided in a cointegration estimation as both the regressand (i.e., output) and regressor (i.e., remittance) are permitted to be in levels, and are non-stationary.

Second, cointegration techniques are designed to address omitted variables and endogeneity issues that generally affect cross-country studies. Notice that remittances rise primarily with

[3]There are a number of studies that do not find a negative effect of remittance on labor supply (see for instance, Vadean, Randazzo, and Piracha, 2019, for the case of Tajikistan). Additionally, studies such as Behzadan and Chisik (2018) also show that remittance can interact with within-country income distribution to generate differences in spending patterns, production patterns, and the pattern of international trade. They find that remittances have the tendency to foster economic growth.

[4]Notice that output and remittance exhibit non-stationary properties; hence, using growth rates or differencing the data means one loses important long run information in the data and only estimates a short run model.

[5]It is important to point out that we do not use remittance as a percentage of GDP in our empirical analysis. We use the natural logarithm of remittances as the regressor to avoid the issues of dividing by GDP as discussed in Clemens and McKenzie (2018).

rising emigration, whose opportunity cost to GDP creates endogeneity bias. Consequently, migration itself becomes an omitted variable correlated with remittances and the error term in the regression, inducing a downward bias of the coefficient on remittance (Clemens and McKenzie, 2018). While cross-country regressions attempt to alleviate or circumvent the endogeneity issues by utilizing instrumental variables in regressions, such strategies may lead to spurious results in the scenario where the instruments are weak or invalid.[6] Moreover, in certain cases it is challenging, and sometimes even impossible to find variables that qualify as valid instruments (Temple, 1999). Cointegration techniques do not require the use of instrumental variables, and by employing them, we circumvent the issue of endogeneity usually present in existing cross-country growth-remittance studies.

Relatedly, as illustrated in (Clemens and McKenzie, 2018), much of the estimated aggregate growth in remittances at the global level during the 1990-2010 period was driven by changes in the measurement rather than the genuine growth in remittance flow. The measurement error issue is likely to lead to non-classical error, which can cloud the true growth-remittance relationship in panel regressions even if remittances are exogenous and the test has high power.[7] Consequently, overstating the growth in remittance flows, then, means that the true effect of that flow on GDP is smaller than would be expected without mis-measurement. However, Hassler and Kuzin (2009) show that even in the presence of measurement errors, cointegration techniques perform better than other competing estimators, making them a desirable estimator in such scenarios.[8]

Beyond addressing the issues discussed in Clemens and McKenzie (2018), we attempt to deal with challenges related to heterogeneity and cross-sectional dependence in panel data head-on. Specifically, existing studies make the stringent assumption of homogeneous coefficient for all cross-sectional units in a panel and hence, assume a common relationship between remittance and output for all the countries in the panel. However, remittances affect output through key observable inputs (e.g., labor and capital) and unobservable inputs such as the multi-factors that affect the level of productivity. Heterogeneity in the relationship between output and remittances is likely to arise due to heterogeneity in technology parameter, indicating differential production function parameters on inputs across countries (see, Eberhardt and Teal, 2010, for a detailed discussion). In our baseline estimation, we employ the between-group dynamic OLS (DOLS) estimator, which allows the remittance-output coefficients to be different across each

[6]See, Herzer and Morrissey (2013) for the case of foreign aid effectiveness.

[7]Non-classical errors occur when there is a potential correlation of the measurement error with the true, unobserved dependent variable itself, with the true values of other variables in the model, or with the errors in measuring those values (Bound et al., 2001).

[8]Miller (2010) also considers a cointegrating regression in which the integrated regressors are messy in the sense that they contain data that may be mismeasured, missing, and observed at mixed frequencies— factors that can generate additional noise that may violate covariance stationarity assumptions on the model error. Miller shows that covariance-based methods for estimating cointegrating regressions may be valid even when the error term is not covariance stationary.

country. The overall remittances-output relationship is therefore interpreted as an average long-run effect of remittances on output in these developing countries.

Furthermore, there are strong arguments that cross-sectional dependence arising from unobserved common factors, economic spillovers, or global shocks is likely to be the norm in panel data rather than the exception (see Baltagi and Pesaran, 2007, for a discussion). More precisely, in a highly integrated global economy, countries interact through among other things trade, immigration, culture and politics. This generates a web of interdependencies within and across countries (Eberhardt and Teal, 2010).[9] As such, assuming cross-sectional independence as done in the standard panel estimations can lead to inaccurate inferences. We attempt to deal with cross-sectional dependence in our baseline estimation by employing the cross-sectionally demeaned data as in Francois and Keinsley (2019) and Herzer and Morrissey (2013), for instance. This strategy can address potential weak cross-sectional dependence. However, to ensure that even in the presence of strong or more complex cross-sectional dependence our baseline findings do not change, we employ the Pesaran (2006) common correlated effects mean group (CCEMG) estimator, and the augmented mean group (AMG) estimator introduced by Eberhardt and Teal (2010).[10] The discussion presented in the above paragraphs indicate that this research offers a robust estimate of the long run remittance-output relationship by making an attempt to address the econometric challenges to detect a positive relationship between the variables while also dealing with the issues related to heterogeneity and cross-sectional dependence in the panel data.

3 Theory: Remittance and Domestic Output in the Long run

The economic theory presented in this section incorporates remittance in a two sector real business open economy model and motivates our empirical specification in Section 4. More importantly, the economic theory sheds light on the need to explicitly model a heterogeneous long run relationship between remittances and output and provides a clear and intuitive interpretation of the coefficient of remittance in the long run relation.

3.1 The Model In our standard neoclassical open economy model the representative household consumes a consumption bundle that consists of a tradable and a non-tradable good denoted by T and N, respectively. The tradable good is the same in the small open economy and in the rest of the world. This small open economy takes as given the world interest rate, given as r^*. For simplicity, we eliminate the government sector.[11]

[9] In the context of our study, assume Ghanaian immigrants in South Africa consistently send money back home to support their families in Ghana. If these immigrants lose their jobs due to an economic downturn in South Africa, this can affect the amount of remittances that they send back home negatively. This in turn can affect economic well-being of these families.

[10] See Eberhardt and Presbitero (2015) for recent application of these estimators.

[11] The model presented in this paper is heavily related to Cardi and Restout (2015), however, we do not focus on fiscal policy— we instead focus on foreign capital flows, particularly, remittances. Additionally, we do not assume

3.2 Household At each instant the representative household consumes a consumption bundle C, which comprises traded goods and non-traded goods denoted by C^T and C^N, respectively. The consumption bundle is defined as follows:

$$C(t) = \left[\varphi^{\frac{1}{\phi}}(C^T)^{\frac{\phi-1}{\phi}} + (1-\varphi)^{\frac{1}{\phi}}(C^N)^{\frac{\phi-1}{\phi}}\right] \tag{3.1}$$

where $\varphi \in (0,1)$ is the weight attached to the traded good in the aggregate consumption bundle and the $\phi > 0$ is the intratemporal elasticity of substitution between tradable and non-tradable consumption goods.

The household agent is endowed with a unit of time and supplies a fraction $L(t)$ of this unit as labor. The agent's leisure is therefore given as $N = 1 - L$. The household derives utility from consumption and disutility from work. The household therefore choose consumption and labor hours worked by maximizing its lifetime utility:

$$\int_0^\infty e^{-\rho t}\left\{\frac{1}{1-\frac{1}{\sigma}}C^{1-\frac{1}{\sigma}} - \chi\frac{1}{1-\frac{1}{\eta}}L^{1-\frac{1}{\eta}}\right\}\mathrm{dt} \tag{3.2}$$

where the parameter $\rho \in (0,1)$ is the discount rate, $\sigma > 0$ is the intertemporal elasticity of substitution of consumption whereas η is the Frisch elasticity of labor supply. The agent maximize their utility subject to the constraints Eqs. (3.3) - (3.4),

$$\dot{B}(t) = r^*B(t) + R(t)^kK(t) + W(t)L(t) - P_C(P(t))C(t) - P(t)I(t) + Rem(t) \tag{3.3}$$

where P_C is the consumption price index, which is a function of the relative price of non-traded goods P. P_CC and PI are the the representative household's expenditure on consumption goods and investment expenditure, respectively. Factor income are derived by supplying labor L and capital K at a wage rate W and rental rate R^k, respectively. Moreover, the agents has access to international borrowing; hence, the accumulate internationally traded bonds, $B(t)$, that yield net interest rate earnings of $r^*B(t)$. The term Rem is the external inflow of remittance to the household agent. Furthermore, capital accummulation follows the dynamic equation

$$\dot{K}(t) = I(t) - \delta K(t) \tag{3.4}$$

The setup above is a standard intertemporal optimization problem. Denoting the co-state variable associated with Eq. (3.3) by λ, we obtain the following optimality conditions for the household agent:

endogeneous markups and we assume away the government sector for simplicity. These modifications, in the context of this study, will not change the primary conclusions of the paper but will induce additional mathematical verbage. Extension of these precluded features is however straightforward.

$$C^{\frac{1}{\sigma}} = P_C\lambda \tag{3.5}$$

$$\chi L^{\frac{1}{\eta}} = \lambda W \tag{3.6}$$

$$\dot{\lambda} = (\rho - r^*)\lambda \tag{3.7}$$

$$r^* = \frac{R^k}{P} - \delta + \frac{\dot{P}}{P}, \tag{3.8}$$

and the relevant transversality conditions

$$\lim_{t\to\infty} \lambda B \mathrm{e}^{-\rho t} = \lim_{t\to\infty} \lambda K \mathrm{e}^{-\rho t} = 0, \tag{3.9}$$

which ensure that household's budget constraint is met.

Furthermore, it is well-known that in an open-economy model with a constant rate of time preference, perfect access to world captial markets, as well as, a representative agent with perfect foresight, a stationary consumption path requires ρ and $r*$ to equal (i.e., $\rho = r^*$). This assumption implies that the marginal utility (shadow value) of wealth λ, will undergo a discrete jump when the agent receives new information and must remain constant over time thereafter (i.e., $\lambda = \bar{\lambda}$).

It is important to note that the homogeneity of $C(.)$ permits for a two-stage consumption decision. Specifically, in the first stage, consumption is determined, and the intratemporal allocation between traded and non-traded. Following, Cardi and Restout (2015) and applying Shephard's lemma, we obtain $C^N = P'_C C$ where P'_C is the partial derivative of P_C with respect to P. Moreover, denoting $\alpha_C = \frac{(1-\varphi))P^{1-\phi}}{\varphi+(1-\varphi)P^{1-\phi}}$ as the share of non-traded goods in the consumption expenditure, the following relations emerge: $C^N = \alpha_C \frac{P_C}{P}$ and $C^T = (1-\alpha_C)P_C C$.[12]

3.3 Firms

There is a representative that produces the tradable good and another firm producing the non-tradable good. These firms act in perfect competition. The representative traded and non-traded sectors utilize, respectively, physical capital, K^T and K^N and labor L^T and L^N to produce traded and non-traded good according to the Cobb Douglas production functions, $Y^T = (K^T)^{\theta^T}(L^T)^{1-\theta^T}$ and $Y^N = (K^N)^{\theta^N}(L^N)^{1-\theta^N}$. The parameter θ^T and θ^N represent the capital income share in output in traded and non-traded sector, respectively. Capital and labor supplied by the household agent are mobile across sectors; hence, they should cost the same price. Consequently, maximizing profits from each sector, the marginal products in the traded and the non-traded sector equalize to:

$$\theta^T(k^T)^{\theta^T-1} = P\theta^N(k^N)^{\theta^N-1} \equiv R^k \tag{3.10}$$

[12]Details on the derivation of these expressions are available in the Appendix.

$$(1-\theta^T)(k^T)^{\theta^T-1} = P(1-\theta^N)(k^N)^{\theta^N-1} \equiv W \tag{3.11}$$

Define $f_k(k^T) = \theta^T(k^T)^{\theta^T-1}$ and $h_k(k^N) = P\theta^N(k^N)^{\theta^N-1}$, where $k^N = K^N/L^N$ and $k^T = K^T/L^T$.

Finally, the following resource constraint arise after aggregating labor and capital over the two sectors

$$L = L^N + L^T \tag{3.12}$$

$$K = K^N + K^T \tag{3.13}$$

3.4 SHORT-RUN EQUILIBRIUM By expressing Eqs.(3.10) and (3.11) in intensive form, we can solve for the sectoral capital-labor ratios as a function of P where $k^T = k^T(P)$ and $k^N = k^N(P)$. It is well-known from two-sector trade models that the relationship between price P and k^T andk^N depends crucially on the sectoral capital intensities (e.g., Brock and Turnovsky, 1994; Cardi and Restout, 2015).[13]. For instance, an increase in the realtive price of non-tradable goods, P will induce resources to shift from the traded to nontraded sector. Now, if the nontraded sector is more capital intensive, capital increase in relative scarcity, which causes the wage-rental ratio to fall and hence, generating the substitution of labor for capital in both sectors. Additionally from Eq.(3.11) the wage rate W is by definition a function of P (i.e., $W = W(P)$).

Given $W = W(P)$, we can substitute into Eq. (3.6) and solve the system in (3.5)-(3.6) to arrive at $C = C(\bar{\lambda}, P)$ and $L = L(\bar{\lambda}, P)$. As we show in the appendix, $C_{\bar{\lambda}} < 0$, $C_P < 0$, $L_{\bar{\lambda}} < 0$, and $L_P \lesseqgtr 0$. Here, a rise in the shadow value of wealth induces the household agent to reduce real expenditure and raise their labor supply. Moreover, an increase in P leads to an appreciation in the relative price of non tradables, which in turn reduces consumption. However, the response of an increase in P depends on which sector is more capital intensive. More precisely, when $k^T > k^N$ ($k^N > k^T$) , an increase in the consumption price index raises (reduces) the wage rate, W[14], which stimulates (depresses) labor supply.

Substituting the sectoral capital-labor ratios into the resource constraints and the production function yields to the short-run static solution for sectoral output where $Y^N = Y^N(K, L, P)$. As explained in Cardi and Restout (2015), according to Rybczynki effect, an increase in capital raises the output of the sector which is more capital intensive. On the other hand, a rise in L will increase output of the sector which is more labor intensive.[15] Finally, an increase in the relative prices of non-tradables exerts opposite effects o sectoral outputs by moving resources away from the traded sector towards the non-traded sector.

[13] $k_P^T = \frac{\partial k^T}{\partial P} = \frac{h}{f_{kk}(k^T-k^N)}$ and $k_P^N = \frac{\partial k^N}{\partial P} = \frac{f}{P^2 h_{kk}(k^N-k^T)}$ We present mathematical details in the Appendix

[14] $W_P = -\frac{k^T h}{k^N-k^T}$

[15] $Y_K^T = \frac{f}{k^T-k^N}$ and $Y_K^N = -\frac{h}{k^T-k^N}$

3.5 Equilibrium Dynamics

The accumulation equation equation for physical capital clears the non-traded good market along the transitional path below

$$\dot{K} = Y^N(K, P, \bar{\lambda}) - C^N(P, \bar{\lambda}) - \delta K \tag{3.14}$$

The adjustment of the open economy toward the steady-state can therefore be described by the dynamic system, which comprises Eq. (3.14) and Eq. (3.6). It is clear from Eq. (3.6) that the price of non-traded goods equalizes the return on domestic capital and traded bonds, r^*. Linearizing the system Eq. (3.6) – (3.14) about the steady state (where the steady state of a variable,X is denoted by $\bar{X}$, the dynamics of K and P can be approximately by

$$\begin{pmatrix} \dot{K} \\ \dot{P} \end{pmatrix} = \begin{pmatrix} a_{11} & a_{12} \\ 0 & a_{22} \end{pmatrix} \begin{pmatrix} K - \bar{K} \\ P - \bar{P} \end{pmatrix} \tag{3.15}$$

where $a_{11} = Y_K^N - \delta$; $a_{12} = Y_P^N - C_P^N$ and $a_{22} = \frac{Y_K^T}{P}$. With the assumption that the system in Eq. (3.15) has one negative eigenvalue denoted by ν_1 and one positive eigenvalue denoted by ν_2, the solution to the system is given as[16]:

$$K - \bar{K} = B_1 e^{\nu_1 t} + B_2 e^{\nu_2 t} \tag{3.16}$$

$$P - \bar{P} = \omega_2^1 B_1 e^{\nu_1 t} + \omega_2^1 B_2 e^{\nu_2 t} \tag{3.17}$$

where B_1 and B_2 are constants to determined. Meanwhile, ω_2^1 and ω_2^2 are the elements of the eigenvector associatd with the eigenvalue ν_1 and ν_2, respectively. While capital stock evolves gradually, the relative price may jump in response to new informations. Consequently, to rule out unstable paths we have set $\omega_2^1 = 0$ (see, Cardi and Restout, 2015, for details).

Finally, setting the current account (CA) to the evolution of foreign asset $\dot{B}$ and substituting Eq. (3.14) into Eq. (3.3), the dynamic equation for the current account equation is given as

$$\dot{B} = r^* B + Y^T(K, L, P) - C^T(P, \bar{\lambda}) + Rem \tag{3.18}$$

Now, linearizing Eq. (3.18) around the steady-state and substituting in Eq. (3.14), we obtain the solution for the stock of foreign asset equation as

$$B - \bar{B} = [(B_0 - \bar{B}) - \psi_1 B_1 - \psi_2 B_2] e^{r^* t} + \psi_1 B_1^{\nu_1 t} + \psi_2 B_2^{\nu_2 t} \tag{3.19}$$

[16]This is can be verified and we present the results in the technical Appendix.

3.5.1 Stationary state This section uncovers the long run impact of a permanent change in remittance (Rem) on the key macroeconomic variables that in turn drives GDP. To help with ease of discussion we present summaries of the main equations. We relegate the mathematical details of the model to a technical appendix. Specifically, following the short-run static solution, the steady-state of the described economy is derived by setting $\dot{B} = \dot{P} = \dot{K} = 0$. The steady-state equilibrium is given by the following equations:

$$h_{kk}(k^N(\bar{P})) = r^* + \delta \tag{3.20}$$

$$Y^N(\bar{K}, \bar{P}, \bar{\lambda}) - C^N(\bar{P}, \bar{\lambda}) - \delta \bar{K} = 0 \tag{3.21}$$

$$r^*\bar{B} + Y^T(\bar{K}, \bar{P}, \bar{\lambda}) - C^T(\bar{P}, \bar{\lambda}) = -\overline{Rem} \tag{3.22}$$

$$B_0 - \bar{B} = \psi_1(K_0 - K) \tag{3.23}$$

We can further substitute Eq. (3.23) into Eq. (3.22) and totally differentiate the resulting system. Now, solving for $\mathrm{d}P$, $\mathrm{d}\lambda$, and $\mathrm{d}K$, we can derive the long run impact of remittance on the following variables

$$\frac{\mathrm{d}\bar{P}}{\mathrm{d}Rem} = 0, \;\; \frac{\mathrm{d}\bar{K}}{\mathrm{d}Rem} = \frac{h_{kk}k_p^N}{\Delta}(Y_\lambda^N - C_\lambda^N) \lesseqgtr 0, \;\; \frac{\mathrm{d}\bar{\lambda}}{\mathrm{d}Rem} = -\frac{h_{kk}k_p^N}{\Delta}(Y_K^N - \delta) < 0, \;\text{and}\; \frac{\mathrm{d}\bar{B}}{\mathrm{d}Rem} = \psi_1 \frac{\mathrm{d}\bar{K}}{\mathrm{d}Rem} \lesseqgtr 0 \tag{3.24}$$

where $\Delta = h_{kk}k_p^N\{(Y_K^N - \delta)(Y_\lambda^T - C_\lambda^T) - (Y_K^T + \psi_1 r^*)(Y_\lambda^N - C_\lambda^N)\}$ is the determinant of the system. Additionally, with the expressions in Eq. (B.57) we can derive the long run impact of remittance on consumption and investment, which can assist in our discussion that follow in section 2.3. The steady state relationships are provided below':

$$\frac{\mathrm{d}\bar{C}}{\mathrm{d}Rem} = -\sigma C \frac{\mathrm{d}\bar{\lambda}}{\mathrm{d}Rem} > 0 \tag{3.25}$$

$$\frac{\mathrm{d}\bar{I}}{\mathrm{d}Rem} = \delta \frac{h_{kk}k_p^N}{\Delta}(Y_\lambda^T - C_\lambda^N) \lesseqgtr 0 \tag{3.26}$$

3.6 Comparative statistics: A permanent increase in remittances We now turn our attention to explaining the behavior of the described economy to a permanent increase in remittances (Rem). The long run impact of an increase in remittances can be inferred from the expressions Eq.(16)–Eq. (18) in the previous section. Here, a permanent increase remittances has no effect on the relative price of nontraded goods, $\bar{p}$ (i.e., $\mathrm{d}P/\mathrm{d}\overline{Rem} = 0$). Consequently, since the long run capital intensities, k^N and k^T, are solely a function of p, the increase in remittances has no impact on the sectoral capital intensities. The steady state increase in remittances however raises the long-run wealth of the economy. With the risk averse behaviour of the households, this lowers the shadow value of wealth (i.e., $\mathrm{d}\bar{\lambda}/\mathrm{d}\overline{Rem} < 0$). With an increase in wealth and constant prices, the economy will realize a rise in the steady state consumption of traded and nontraded goods. Hence, the increase in remittance in the modelled economy will unambiguously raise aggregate consumption. The increase in demand for consumption of nontraded goods will induce a rise in the output of nontraded goods, which will in turn attract labor from the traded sector of the economy causing a decline in output in the traded sector.

As previously discussed, the sectoral intensities are not affected following the increase in remittances, the long run impact of remittance on capital (K) is ambiguous (see Eq. (B.57)) and critically depends on whether labor is moving from a relatively less capital intensive to a relatively more capital intensive sector. In this scenario, if the nontraded sector is relatively intensive in its use of capital ($k^N > k^T$), remittance will have a positive permanent effect on capital. The dynamics also holds for steady state investment since it is proportional to capital stock in the steady state, governed by depreciation. Now,

$$\overline{gdp} + \overline{Rem} = \bar{C} + \bar{I} + \overline{TB} \tag{3.27}$$

Totally differentiating Eq. (3.27) and dividing by $\mathrm{d}Rem \neq 0$ and solving for gdp in terms of model parameters and remittances, we obtain

$$\frac{\mathrm{d}\overline{gdp}}{\mathrm{d}\overline{Rem}} = \underbrace{\left(\sigma\frac{h_{kk}k_p^N}{\Delta}(Y_K^N - \delta)C + \delta\frac{h_{kk}k_p^N}{\Delta}(Y_\lambda^T - C_\lambda^N) + r\psi\frac{h_{kk}k_p^N}{\Delta}(Y_\lambda^N - C_\lambda^N) - 1\right) \lesseqqgtr 0}_{\text{Composite long-run effect of remittance on output}} \tag{3.28}$$

Main Takeaway The theory in the previous section highlights two important points that motivates the empirical work that follow. First, it is clear from the theory that the impact of remittance on output is ambiguous due to the potential opposing effects that can arise from the labor, investment, and consumption channels through which remittance can impact GDP. The implication for the empirical work is that if one controls explicitly for investment for instance in a remittance–output regression, then one would be precluding a potential channel through which remittance can affect output. Hence, working in a bivariate framework allows one to investigate all the potential channels for which remittance can impact real GDP. Second, the ambiguity in the relationship between remittance and the key macro-aggregates, which is dependent on the parameterization of the model, suggests that an assumption of a homogeneous relationship between output and remittance across countries is stringent. More precisely, one would expect a more realistic assumption of a heterogeneous relationship across countries. To this end, we proceed to estimate a heterogeneous bivariate cointegration model of output and remittances.

4 Evidence

The primary equation for estimation is given as

$$Y_{it} = \alpha_i + \beta Rem_{it} + \delta_i t + \varepsilon_{it} \quad (4.1)$$

where Y_{it} is defined as the natural logarithm of real GDP in country $i = 1, 2, \ldots, N$ at time t. Rem_{it} is the natural logarithm of real remittance. The parameter of interest β captures the average impact of remittance on output. The remittance-output coefficient is estimated for each country, and is heterogeneous across countries in the panel estimate. We discuss this in detail in section 5. The parameters a_i and δ_i represent the constant and linear time trends associated with each country i, respectively. Finally, we assume the error term ε_{it} to be stationary. This assumption is important because if we have omitted a non-stationary variable, which forms part of the potential cointegration relation in Eq. (4.1), then the error term will exhibit non-stationary properties in which case cointegration in Eq. (4.1) would fail to hold. We present formal residual-based cointegration tests for this stationarity assumption in the section 4.1.2.

4.1 Data and Pre-testing

The dependent variable is real output and it is measured by GDP (constant 2010 US$). The explanatory variable, real remittance inflows is measured as personal remittance inflows (constant 2010 US$) and is computed by deflating the nominal migrant remittances by GDP deflator. Annual data on GDP and remittance come from the World Development Indicators 2019 database.[17] A well-known challenge with utilizing remittance data is the issue of uneven reporting across countries, which imposes large missing values in the dataset (Abdih et al., 2012). While we start our analysis with 137 developing countries, several of the countries were missing data for many periods. In fact, preliminary organization of the data shows that over the period 1970-2017, only four countries, Algeria, Colombia, the Dominican Republic and South Africa, had complete data on remittances. Moreover, allowing for balanced data, which will be ideal, induces a tradeoff in the number of time periods and countries that we can include in the data. A compromise here is therefore to employ an unbalanced panel data, which in itself poses some challenges. Specifically, the number of periods T required to carry out the pre-testing exercises require T to be greater than or equal to a minimum value for a given number of lags. For instance, with 3 lags as well as constant and trend terms, at least 12 observations are required for each cross-sectional unit to successfully conduct the Pesaran (2007) unit root test. Finally, while the unit roots tests employed in the paper can be applied to unbalanced panel data, they do not permit for gaps in the data.[18] Addressing all of these

[17] Data on remittances is available at the World Bank's Migration and Remittances Data here: https://www.worldbank.org/en/topic/migrationremittancesdiasporaissues/brief/migration-remittances-data. We employed the most recent version of the data, which at the time of writing this paper was the Annual Remittance Data, updated as of December 2018.

[18] See Martins (2011) for a neat summary of the characteristics of unit root tests with respect to unbalanced data.

issues with respect to the data, we arrive at an unbalanced panel (without gaps) with a sample of 80 countries representing approximately 58% of all developing countries for the period 1970-2014.[19].

At this point, we focus on the time series properties of our data. Specifically, since our focus is on the long run impact of remittance on output, steady state (permanent) changes in remittance are associated with permanent changes in real GDP. Consequently, we test for: (a) the presence of unit root in the individual time series, $\ln(Y_{it})$ and $\ln(Rem_{it})$, and (b) the existence of a cointegration relation between $\ln(Y_{it})$ and $\ln(Rem_{it})$.

4.1.1 Unit Root Tests

To test for the presence of unit root in the data, we employ the panel unit root test of Im, Pesaran, and Shin (2003, henceforth IPS), which is based on the Augmented Dickey-Fuller (ADF) regression for the individual cross-section unit in the panel. However, given the likelihood of common shocks or spillovers across countries, the error terms ε_{it} may not be independent. In the presence of a potential cross-sectional dependence, the IPS test can lead to spurious inferences. We therefore consider the cross-sectionally augmented IPS (CIPS) test proposed by Pesaran (2007). The CIPS filters out any cross-section dependency by augmenting the ADF regression with the cross-section averages of lagged levels and first-differences of the individual series (See for example Baltagi and Pesaran, 2007; Herzer and Grimm, 2012, for a discussion on second generation unit root tests).

Table 1: Panel Unit Root Tests

	Deterministic trend	IPS statistics	CIPS statistics (Zt-bar)
Level Data			
$\ln(Y_{it})$	c, t	0.0703	4.138
$\ln(Rem_{it})$	c, t	-1.4057	1.065
First-Differenced Data			
$\Delta \ln(Y_{it})$	c	-10.0676^{***}	-3.249^{***}
$\Delta \ln(Rem_{it})$	c	-12.5166^{***}	-8.205^{***}

Notes: *** indicates significance at the 1 percent level. For the level data, we allow for both individual country effects (c) and country-specific time trends (t). In the case of the first-differenced data, we allow for individual country effects (c). Four lags were selected to adjust for autocorrelation. The IPS statistic is distributed as $N(0,1)$. For CIPS under unbalanced panel, only standardized Zt-bar statistic can be calculated. The CIPS statistic assumes cross-sectional dependence in data.

As can be seen from Table 1, both the IPS and CIPS fail to reject the unit-root null hypothesis for the level data, however, they strongly reject the null for the first differenced series. These

[19]There is growth in the use of cointegration techniques for unbalanced panels. See, Rajbhandari and Zhang (2018) and Eberhardt and Teal (2019) for DOLS and FMOLS applications, respectively. The list of countries and additional information on the data are available in Appendix C.Additionally, we report the country-by-country summary statistics for GDP and remittances in Appendix D

findings suggest that the individual series in Eq.(4.1) are non-stationary I(1) processes.

4.1.2 Cointegration Tests

As shown in section 4, our simple model proposes a potential long-run relationship between remittances and output. Thus, before proceeding to estimate the long-run impact of remittances on output it is important to check if there exists an empirical long-run relationship between the two variables. To do this, we test for the presence of cointegration in Eq. (4.1). We employ four standard panel and group test statistics suggested by Pedroni (1999). The standard Pedroni tests, however, do not account for potential cross-sectional dependence. In the presence of cross-sectional dependence that may arise from multiple unobserved common factors, an assumption of cross-sectional independence can lead to biased inference (Baltagi and Pesaran, 2007; Herzer and Morrissey, 2013). To account for cross-sectional dependence, we utilize the four standard Pedroni test but include common time dummies to address cross-sectional dependency in the manner of Neal (2014). This strategy involves time demeaning of the data for each cross-sectional unit and variable (See, Neal, 2014, for the theory and implementation details). To highlight the relevance of accounting for cross-sectional dependence in the data, we report test results in which we assume cross sectional independence.

As the Pedroni tests are residual based, they allow to formally test whether our stationarity assumption of the error term in Eq. (4.1) indeed holds. The null hypothesis of the Pedroni tests is no cointegration, meaning the error term is not stationary. This also implies that if the test results reveal a failure to reject the null hypothesis, then our specification of a bivariate relationship between output and remittances is incorrect. This may be due to an omission of a non-stationary variable, which should be part of the cointegration system. On the other hand, a rejection of the null hypothesis of no cointegration for the residual based tests is a validation of our simple bivariate specification, which also suggests that issues of omitted variables and measurement errors, to a large extent, are not problematic.

To complement the residual-based tests, we utilize the approach of Larsson, Lyhagen, and Löthgren (2001), which is based on Johansen's (1988) maximum likelihood estimation procedure to test for cointegration among the variables in Eq. (4.1). The test allows us to confirm the presence of one cointegrating vector in Eq. (4.1). It is important to note that although the Larsson et al. panel test does not account for potential cross-sectional dependence in the data, it treats all variables as potentially endogenous. It therefore avoids the normalization issues inherent in residual-based cointegration tests. The null hypothesis for Larsson et al. panel test is that all countries have the same number of cointegrating vectors denoted by r_i among the p variables, in our case two variables as in Eq. (4.1). More precisely, $H_0 : rank(\Pi_i) = r_i \leq r$ and the alternative hypothesis is $H_1 : rank(\Pi_i) = p$, for all $i = 1, ..., N$ where Π_i is the long-run matrix of order $p \times p$. For completeness we also report the Fisher statistic proposed by Maddala and Wu (1999). [20]

Table 2: Panel Cointegration Tests

	Panel A: Pedroni Residual-Based Tests	
Statistics	Cross-sectional dependence	Cross-sectional independence
Panel PP statistic	-2.750^{***}	-0.438
Panel ADF statistic	-2.043^{***}	-1.259
Group PP statistic	-3.454^{***}	-0.729
Group ADF statistic	-2.818^{***}	-2.053^{***}
	Panel B: Cointegration Rank Tests	
	$r = 0$	$r = 1$
Panel standardized trace statistic	17.60^{***}	0.621
Fisher statistic	787.9^{***}	166.2

Notes: *** denotes a rejection of the null hypothesis of no cointegration at the 1% level. ADF stands for *augmented Dickey Fuller* and PP stands for *Phillips-Perron*. For the Pedroni tests the number of lags was determined by the Schwarz criterion with a maximum of three lags. All test statistics are distributed N(0,1), under a null hypothesis of no cointegration. For the cointegration rank test the Schwarz criterion suggests two lags.

Table 2 presents the test results. In Panel A, it is clear that under the assumption of cross-sectional dependence all the tests decisively reject the null hypothesis of no cointegration at the 5 percent level or better. In contrast, only the Group ADF statistic rejects the null of no cointegration when cross sectional independence is assumed. This highlights the relevance of explicitly accounting for cross sectional dependence, which is often the rule rather than the exception in panel data analysis. Moreover, as shown in Panel B, both the standardized trace and Fisher statistics support the presence of one cointegration vector in the case of Eq. (4.1). Together these results validate our bivariate empirical specification.

5 Estimation and Results

We now turn our focus to uncovering the long-run impact of remittances on output. Following the strong evidence of unit roots in the two variables of interest, and cointegration between output and remittances, we can consistently estimate the long-run relationship in Eq. (4.1) using recently developed cointegration techniques. We employ the between-dimension group-mean panel dynamic ordinary least squares (DOLS) estimator of Pedroni (2001b). The panel

[20] Computational details as well as theory related to this test are laid out in Larsson et al. (2001) and in the appendix of Herzer and Morrissey (2013) for example.

DOLS regression is given by[21]

$$Y_{it} = \beta_i Rem_{it} + \sum_{j=-p_i}^{p_i} \Psi_{1ij} \Delta Rem_{it-j} + \alpha_i + \delta_i t + \varepsilon_{it}, \tag{5.1}$$

where Ψ_{1ij} are coefficients of lead and lag differences which account for potential serial correlation and endogeneity of the regressors. The DOLS estimator produces unbiased estimates for variables that are cointegrated even in the presence of endogenous regressors. This feature is important in our case because as previously discussed, remittances are more likely to be endogenous than exogenous. Furthermore, the group-mean panel DOLS estimator is super consistent under cointegration and robust to any omitted variables suggesting that the use of instrumental variables to address any omitted variable problem is not required. Recall that the parameter of interest is β and under the group-mean DOLS estimator is computed as

$$\hat{\beta} = \frac{1}{N} \sum_{i=1}^{N} \hat{\beta}_i, \tag{5.2}$$

where $\hat{\beta}_i$ is the conventional time-series DOLS estimator applied to the ith cross sectional unit (i.e., country) of the panel.

Recently, there has been strong arguments that interdependencies due to common shocks or global spillovers among countries at the same time are likely to be the norm in panel data (Baltagi and Pesaran, 2007; Neal, 2014). Indeed, our results from the cointegration tests suggest that accounting for these potential cross-sectional dependence is required. While there are a number of ways to account for these cross- sectional-dependencies, we apply the DOLS procedure to demeaned data in the manner of Herzer and Morrissey (2013) and Francois and Keinsley (2019), among others.[22] More precisely, in place of Y_{it} and Rem_{it} in the baseline equation in Eq.

[21]It is important to note that an alternative estimator to DOLS is the fully-modified OLS (FMOLS) estimator proposed by Pedroni (2001a). As Pedroni (2001a) discusses, the FMOLS requires fewer assumptions and tends to be more robust. Moreover, because DOLS involves adding lead and lagged difference terms of the independent variables, in cases in which there are several independent variables and limited time observations, the DOLS estimator can result in severe loss in degrees of freedom (Liddle, 2012). However, our specification is parsimonuous and is unlikely to suffer from the latter. More importantly, Kao and Chiang (2000) illustrate that for the case of a single regressor, the DOLS estimator has a smaller bias than FMOLS. Consequently, the DOLS is our preferred estimator, however, we also report the results from FMOLS estimation.

[22]This strategy of demeaning the data avoids overparamerterization as it preserves the number of regressors in the specification. However, alternative strategies such as the Pesaran (2006) common correlated effects (CCE) approach where observed regressors are augmented by cross-sectional averages of the dependent variable and the individual specific regressors increase the number of regressors. Specifically, if we utilize this approach, we have to include $\overline{Y}_t$ and $\overline{Rem}_t$ and their respective leads and lags in our specification, which will increase the number of parameters to estimate and likely weaken the power of the DOLS estimator due to overparameterization. Although we do not report the results in the paper, the estimated β using this augmentation option to control for cross sectional dependence is 0.045 and 0.014 for the DOLS and FMOLS, respectively. We also present results from the Pesaran (2006) CCE and augmented mean group by Eberhardt (2012) estimator by in section 6.5.

(5.1), we utilize $\widetilde{Y}_{it}$ and $\widetilde{Rem}_{it}$ where

$$\widetilde{Y}_{it} = Y_{it} - \overline{Y}_t, \qquad \text{where } \overline{Y}_t = \frac{1}{N}\sum_{i=1}^{N} Y_{it},$$

and

$$\widetilde{Rem}_{it} = Rem_{it} - \overline{Rem}_t, \qquad \text{where } \overline{Rem}_t = \frac{1}{N}\sum_{i=1}^{N} Rem_{it},$$

For completeness, we report the results for both the data in which cross-sectional dependence and independence are assumed. However, we place more weight on the results from the model in which we account for cross sectional dependence as it is likely to be the rule in panel data. Additionally, we report results from the fully modified OLS estimators.

Table 3: Estimated long-run relationship between remittances and output

	Data Treatment	
Estimator	C-S Dependence	C-S Independence
DOLS	0.0664***	0.0407***
	(6.0365)	(4.3050)
FMOLS	0.0638***	0.0277***
	(8.1009)	(4.2465)

Notes: The dependent variable is the natural log of GDP. *** indicates significance at the 1 percent level. C-S is cross-sectional. *t*-statistics in parentheses. The number of leads and lags in the individual DOLS regressions was fixed to 1 lag. The model allows for individual and time fixed effects.

Table 3 reports the baseline results from the cointegration regressions. The coefficient of remittance β is consistently positive, statistically significant, and more importantly, economically large across all four regressions in our baseline case. It is however clear that estimates of β from the demeaned data, which accounts for common shocks and spillovers, are generally larger than the results from the data in which cross sectional independence is assumed. As mentioned earlier, we place more emphasis on the results from the demeaned data. Consequently, the result from the DOLS regression suggests that a 1 percent increase in remittances will, on average, generate an increase of 0.066 percent in real GDP in the long run. Recall from the theoretical model that β captures the total effect of remittances on output; hence, we can interpret this observed positive effect as the overall long-run impact of remittances on output. From the economic sense, this means that in the long-run, the positive effects of remittances on output through the channels discussed in section 3 outweighs any potential negative effect. Our result is in sharp contrast to the studies such as Barajas et al. (2009) who find that, at best, workers' remittances have no impact on economic growth and Jahjah et al. (2003), who find negative effect of workers remittance on growth. Moreover, Herzer and Morrissey (2013) find a negative long-run relationship between foreign aid and output.[23] The results in this paper therefore suggest that remittances are not simply outgrowing aid in terms of size, but they may be an important driver of long-term output than foreign aid in developing countries.

6 Robustness

In this section, we conduct a series of robustness exercises to ensure that our baseline positive remittance-output estimate is consistent across income groups, and periods and not driven by potential outliers, sampling bias, or model specification.

6.1 Income Groups A natural question that arises in our study is: Is the long run impact of remittance on output different across income groups? The rationale here is that countries in different income groups have differential characteristics, and more importantly, are at different stages of development. It is therefore important to investigate whether the positive long-run remittances-output effect is not driven by a particular income group; hence, development level. We split our sample into two income groups, low- and lower-middle income and upper-middle income groups resulting in 47 low and lower-middle income and 33 upper-middle income countries. We re-estimate the DOLS and FMOLS regressions for these two income groups. In both cases, we account for potential cross-sectional dependence in the data.

Table 4: Estimated long-run relationship between remittances and output by income level

	Income Group	
Estimator	Upper-Middle Income	Low and Lower-Middle Income
DOLS	0.1167***	0.0759***
	(4.5904)	(8.1487)
FMOLS	0.1067***	0.0833***
	(6.0334)	(9.9746)

Notes: The dependent variable is the natural log of GDP. *** indicates significance at the 1 percent level. Upper middle income countries consists of 33 countries. Low- and lower-middle income countries comprises 47. *t*-statistics are in parentheses. The number of leads and lags in the individual DOLS regressions was fixed to 1 lag. The model allows for individual and time fixed effects. The regressions account for cross-sectional dependence.

Table 4 presents the results. It is clear that regardless of the income group, the effect of remittance on output is positive and statistically significant at the conventional levels. There is however differences in the size of this long run effect. More precisely, the results from the DOLS

[23] We must add that Herzer and Morrissey (2013) find that conditional on aid financing investment, aid has an overall positive effect on output in the long-run. Nonetheless, this conditional positive effect is smaller than our estimated remittance-output impact.

estimation shows that while a 1 percent increase in remittance increases output by 0.12 percent in upper-middle income countries, low and lower-middle income countries miss about 0.044 percentage points of this impact. This suggests that on average, remittances are likely to have a bigger positive long-run effect in upper-middle income countries than in low and lower-middle income countries.

6.2 Sensitivity analysis

We conduct regional and country sensitivity analysis in the manner of Herzer and Morrissey (2013) to ensure that our positive remittance-output relationship is not driven by outliers. This is important because our parameter of interest β is an average of individual coefficient for each country in our sample. Hence, if a single country or group of countries in a particular region have a positive and large estimate of β, then this will likely drive the observed average positive effect of remittance on output. Additionally, country and regional heterogeneity exist in terms of remittance inflows. For instance, remittance inflows ranged from approximately 7 percent of GDP in East Asia and the Pacific to 12 percent in South Asia in 2018. It is therefore important to account for possible outlier effects that may arise from these differences.

Against this backdrop, we begin our sensitivity analysis by re-estimating our baseline DOLS regression by excluding one country at a time from the full sample. Figure 1 shows the estimates of β in the left panel and their corresponding t-statistics in the right panel. Evidently, the estimated long-run effect of remittances on output is consistently positive and ranges between 0.057 and 0.076, values close to the baseline esitmate, 0.066, which is given as the red dashed line in the figure. All of the estimates are statistically significant at the 5 percent significant level or better.

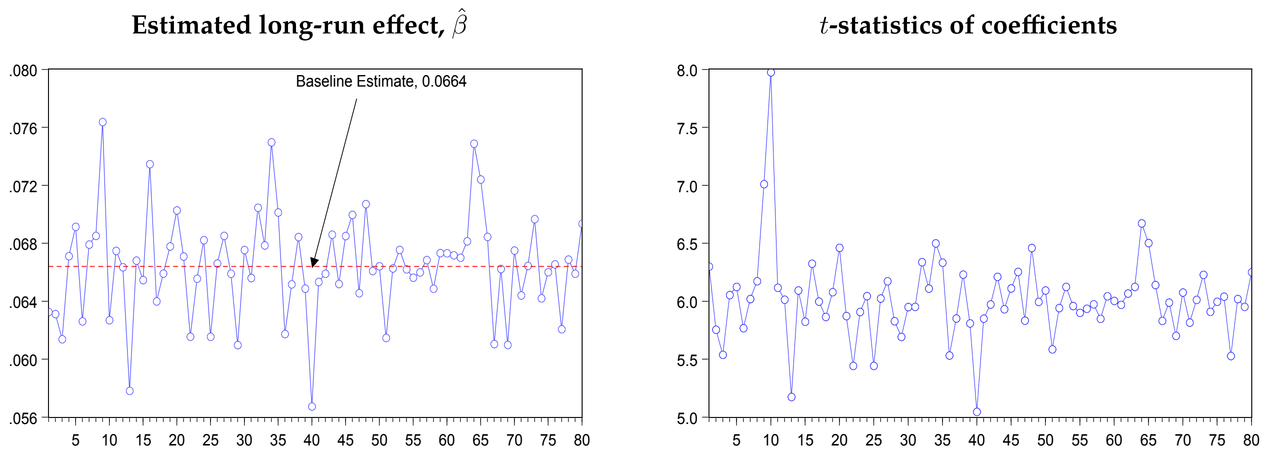


Figure 1: Sensitivity Analysis, DOLS estimation with single country excluded from the sample. The y-axis is the estimated β. The horizontal axis is the ith country left out during the estimation. Specifically, one country is excluded from the panel at each round of estimation. Number of cross-sectional units is therefore 79 for each estimation.

We now turn our attention to conducting a similar sensitivity analysis at the regional level. We employ the DOLS model in which we account for cross sectional dependence in the data. We also report the estimates from the FMOLS estimator. Table 5 presents the results from the regional sensitivity analysis.

Table 5: Estimated long-run relationship between remittances and output, regional sensitivity analysis

Excluded Region	DOLS	FMOLS	No. of countries in Panel
East Asia & Pacific (EAP)	0.0779***	0.0744***	70
	(5.99)	(7.94)	
Europe and Central Asia (ECA)	0.0749***	0.0739***	66
	(6.52)	(8.33)	
Middle East and North Africa (MENA)	0.0576***	0.0559***	72
	(5.11)	(7.08)	
Latin America and Caribbean (LAC)	0.0581***	0.0612***	62
	(5.05)	(6.91)	
South Asia	0.0693***	0.0681***	75
	(6.04)	(8.18)	
Sub-Sahara Africa (SSA)	0.0822***	0.0714***	55
	(5.04)	(6.53)	

Notes: *** indicates significance at the 1 percent level and t-statistics are presented in parentheses. The dependent variable is the demeaned real GDP, $\bar{Y}_{i,t}$. The number of leads and lags in the individual DOLS regressions was fixed to one lag and one lead.

The first column in the table shows the region that was excluded from the estimation. Here, we focus on the result from the DOLS estimates. It is clear that the long-run impact of remittances on real GDP ranges from 0.058 when we preclude MENA or LAC from the sample to about 0.082 when SSA is excluded from the full sample. All of the estimates are statistically significant at the 1 percent level and are in the neighborhood of our baseline estimate of 0.066. The results from both the country and regional sensitivity analysis show that while minor differences may exist in the size of remittance-output effect, the sign of the effect remains consistently positive. These findings strongly suggest that the postive long-run impact of remittance on output is not driven by possible outlier either at the country or regional level. This conclusion holds true for the results from the FMOLS estimations.

6.3 The 1990's and 2000's It is well-known that prior to 1990, remittance data coverage and perhaps, its measurement is poor (See for instance, Clemens and McKenzie, 2018). The 1990s and 2000s were periods of rapid technological advancement in many areas of development. Additionally, these periods witnessed the commercialization of the Internet, which has led to an increased expansion to better and quality data. Hence, one would generally expect better coverage, as well as, measurement of remittance data. In this exercise, we drop the period leading up to 1990 and conduct our empirical analysis for the sub-period 1990 to 2014 in an attempt to capture the potential role of "quality" remittance data. It is however important to note that restricting the analysis to this period comes at a cost as we have fewer time periods, which can impact the detection of a long-run relationship. We acknowledge this potential issue and proceed with the analysis.

Table 6: Estimated long-run relationship between remittances and output, 1990-2014

	Model Assumption for Data	
Estimator	C-S Dependence	C-S Independence
DOLS	0.0199^{*}	0.0269^{***}
	(1.7366)	(2.8570)
FMOLS	0.0217^{***}	0.0170^{***}
	(3.3194)	(2.7832)

Notes: The dependent variable is natural log of GDP. *** indicates statistical significance at the 1 percent level, while * indicates significance at the 10 percent level. C-S is cross-sectional. *t*-statistics are in parentheses. The number of leads and lags in the individual DOLS regressions was fixed to 1 lag and one lead. The model allows for individual and time fixed effects.

Results from the sub-period analysis are presented in Table 6. As can be seen from the table, the effect of remittance on output is consistently positive and significant at conventional levels. Specifically, focusing on the DOLS estimate in which cross sectional dependence is accounted for in the data, we find that a percent increase in remittance will induce a 0.02 percent increase in real output. Although, this effect is non-trivial, the estimated impact from the sub-period analysis is generally smaller than the effect in the baseline scenario. The weaker long-run relationship can be attributed to the likelihood of lower persistence in the series in question as we have used fewer time periods in our long run analysis.

Table 7: Estimated long-run relationship between remittances and output, Per Capita Measure

	Model Assumption for Data	
Estimator	C-S Dependence	C-S Independence
DOLS	0.0481^{***}	0.0362^{***}
	(5.0935)	(3.9566)
FMOLS	0.0402^{***}	0.0218^{***}
	(6.0392)	(3.4028)

Notes: The dependent variable is natural log of per capita GDP. *** indicates significance at the 1 percent level. C-S is cross-sectional. *t*-statistics in parentheses. The number of leads and lags in the individual DOLS regressions was fixed to 1 lag. The model allows for individual and time fixed effects.

6.4 Per capita measures Finally, we employ per capita measures of output and remittances as alternative measures to our baseline measures, and re-estimate Eq. (4.1). Table 7 presents the results. It is clear that remittance per capita has a positive long run impact on GDP per capita. More specifically, the results from the DOLS regression reveal that on average a 1 percent increase in remittance per capita raises GDP per capita by approximately 0.048 percent in the long run.

6.5 Alternative estimators As discussed earlier, cross-sectional dependence arising from global shocks and economic spillovers can lead to inaccurate inference. In section 5, we attempt to account for cross-sectional dependence by employing cross-sectionally demeaned data for the DOLS estimation. However, while the DOLS addresses the primary issues associated with endogeneity, measurement errors, lack of power and heterogeneity, it may not fully address the case where cross-sectional dependence is strong or complex. More precisely, our baseline strategy of cross-sectionally demeaning the data prior to estimation assumes that cross-section correlation is of a nature equivalent to common shocks with identical impact across countries, which is a strong assumption. Here, we relax this assumption and present results from the augmented mean group (AMG) estimator introduced by Eberhardt and Teal (2010) and the Pesaran (2006) common correlated effects mean group (CCEMG) estimator, both of which are designed to address more complex cross-sectional dependence. For instance, the empirical strategy employed by the CCEMG naturally deals with complex cross-sectional dependence and allows for time-variant unobservables with heterogeneous impact across panel members (Eberhardt, 2012). Specifically, the CCEMG addresses identification problems arising from the aforementioned issues with a simple but powerful augmentation of the group-specific regression equation. That is, apart from the standard regressors in our specification in Eq. (4.1), we now include the cross-section averages of the dependent and independent variables, $\bar{Y}$ and $\overline{Rem_{it}}$, respectively, as additional regressors.

There is an important distinction between the CCEMG and the AMG estimator. The CCEMG treats the set of unobservable common factors as a nuisance, with no economic interpretation. The AMG estimator on the other hand captures these unobservables as components of total factor productivity. Hence, in addition to the CCEMG, we employ the AMG estimator, which was developed as an alternative to the CCEMG with production function estimation in mind (Eberhardt and Teal, 2010).

As described in Eberhardt (2012), the AMG procedure is implemented in three steps: First, a pooled regression model augmented with the year dummies is estimated by first-difference ordinary least squares, and the coefficients of the (differenced) year dummies are collected. They represent an estimated cross-group average of the evolution of unobservable TFP over time. This is referred to as "common dynamic process". Second, the group-specific regression model is then augmented with this estimated TFP process: either (a) as an explicit variable or (b) imposed on each group member with unit coefficient by subtracting the estimated process from the dependent variable. To ensure that the positive relationship between output and remittances is not affected by how we treat cross-sectional dependence, we present results from both strategies given in (a) and (b). The regression model includes an intercept, which captures time-invariant fixed effects (TFP levels). Third, similar to the grouped-mean DOLS, FMOLS, and CCEMG estimators, the group-specific model parameters are then averaged across the panel.

To this end, consider the following bivariate model: for $i = 1, ..., N$ (countries) and $t = 1, ..., T$ (years),

$$Y_{it} = \beta_i Rem_{it} + \varepsilon_{it} \tag{6.1}$$

$$\varepsilon_{it} = \alpha_{1i} + \lambda_i f_t + u_{it} \tag{6.2}$$

$$Rem_{it} = \alpha_{2i} + \lambda_i f_t + \gamma_i g_t + \epsilon_{it}, \tag{6.3}$$

where Rem_{it} and Y_{it} are the observables remittance and output, respectively. β_i is the country-specific slope on the observable regressors and ε_{it} contains the unobservables and the error terms u_{it}. The unobservables in Eq. (6.2) are made up of standard group-specific fixed effects α_{1i}, which capture time-invariant heterogeneity across groups, as well as an unobserved common factor f_t with heterogeneous factor loadings λ_i, which can capture time-variant heterogeneity and cross-section dependence. u_{it} and ϵ_{it} are assumed white noise. The factors f_t and g_t are not limited to linear evolution over time; they can be nonlinear and nonstationary. As one can see, this has clear and immediate implications for cointegration. Moreover, problems arise if the regressors are driven by some of the same common factors as the observables. Specifically, the presence of f_t in equations Eq. (6.2) and Eq. (6.3) induces endogeneity in the estimation equation (see, Coakley et al., 2006; Eberhardt and Teal, 2011, for detailed discussion). Both the CCEMG and AMG are designed to deal with the aforementioned issues.

Table 8: Estimated long-run relationship between remittances and output, alternative estimators

Estimator	Unweighted	Weighted
CCEMG	0.0161^{*}	0.0099^{*}
	(0.0088)	(0.0053)
AMG	0.0217^{**}	0.0110^{*}
	(0.009)	(0.0057)

Notes: The dependent variable is the natural logarithm of GDP. ** and * indicate significance at the 5 and 10 percent levels respectively. Standard errors in parentheses.

Table 8 presents the results. The second column does not account for outlier effects in the computation of the mean-group estimator. In contrast, the third column puts less emphasis on outliers while computating the average coefficient from the country-specific estimated coefficients. The results clearly support our baseline conclusion as all estimates are positive and statistically significant at conventional levels.[24]

6.6 Output, Remittance, and Investment

Finally, we control for investment as a major factor of output in the long run relation. By explicitly including investment in the cointegration regression, we are able to isolate the other channels that drives the relationship between remittances and output. More importantly, the results from a cointegration regression of remittance and output including investment allows as draw conditional conclusions on the remittance-investment channel. We employ the cross-sectionally demeaned data for the estimation.

Table 9 presents the results. As shown in the table, a 1 percent increase in remittance is associated with an average increase of 0.048–0.049 percent increase in real GDP in the long run. Interestingly, we observe that the coefficient on investment is positive and statistically significant. This suggests that if the increase in remittances is associated with an increase in investment, this can lead to a stronger positive remittance-output relationship. Additionally, the results implies that other than the investment channel, the long-run positive relationship between remittance and real GDP is mediated by the other channels discussed in the theory in section 3. Finally, the results reinforces the positive remittance-output relationship uncovered in the baseline results.

Table 9: Estimated long-run relationship between remittances and output while accounting for investment

Estimator	DOLS	FMOLS
Rem_{it}	0.0487^{***}	0.0483^{***}
	(5.0762)	(7.6691)
Inv_{it}	0.0124^{***}	0.0102^{***}
	(10.1258)	(13.3372)

Notes: The dependent variable is the natural log of GDP and Inv_t is investment as a share of GDP. *** indicates significance at the 1 percent level.t-statistics in parentheses. The number of leads and lags in the individual DOLS regressions was fixed to 1 lag. The model allows for individual and time fixed effects and accounts for cross-sectional dependence via cross-sectional demeaning. The Fisher statistics support the presence of one cointegration vector in this specification.

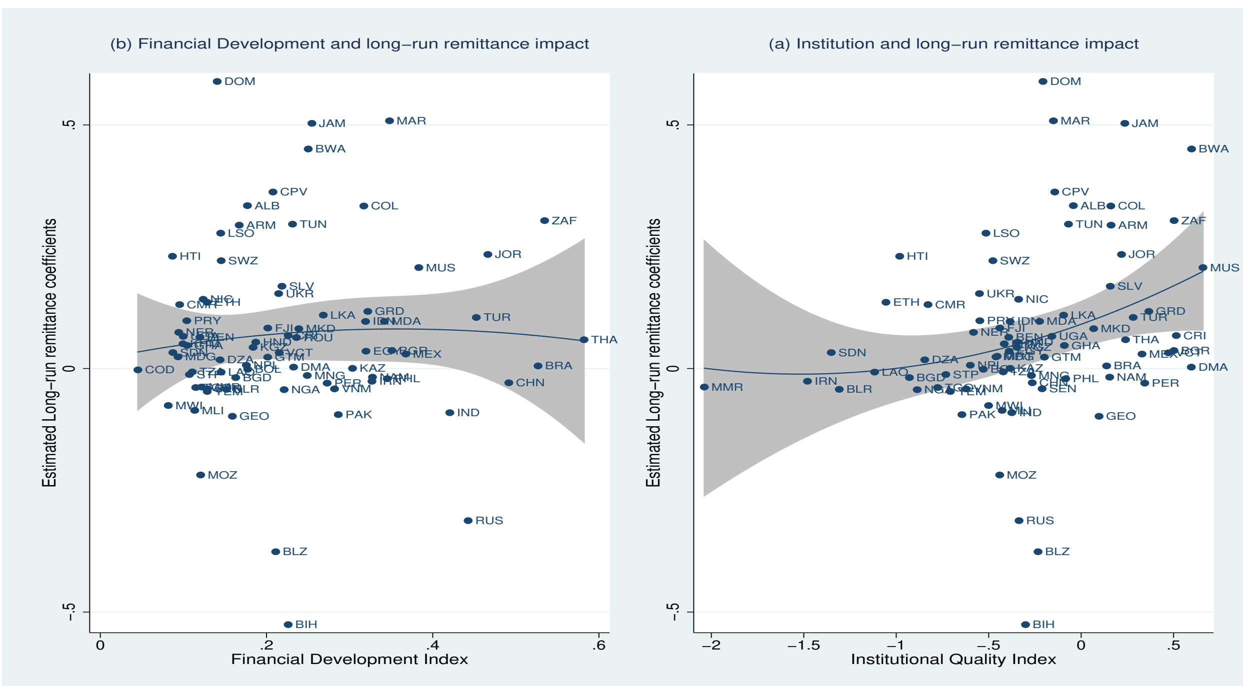
[24]In addition to these estimators, we further employ recently developed estimators, the distributed lag and Autoregressive distributed lag approach by Chudik et al. (2017). The results are comparable to the other estimators employed here. Results are available upon request.

7 Explaining the variation in the long run impact of remittance

Recall that the theory in section 2 motivated a need to estimate a heterogeneous remittance-output relationship. The empirical estimations in the previous sections do allow for variations in the country-specific relationship between output and remittances. This section sheds light on the plausible explanations for the heterogeneous relationship between remittance and output across countries. To this end, we employ two important and well-known growth-facilitating variables— financial development and institutional quality— to explain the cross-section *dispersion* of the long run remittance coefficient. [25].

We collect the estimated remittance coefficients in the baseline estimation from the previous section and investigate how they relate to financial development and institutional quality in the countries under study.

Figure 2: Patterns for baseline DOLS remittance coefficients



Notes: We plot the country-by-country long-run coefficients for remittances. We use the estimated coefficients from the DOLS estimator that accounts for cross sectional dependence in the data. The data on Financial Development is from the International Monetary Fund and it available at: https://data.imf.org/?sk=F8032E80-B36C-43B1-AC26-493C5B1CD33B. Data on the measure of Institutional Quality proxied by regulatory quality is from World Governance Indicators constructed by Kaufmann et al. (2011). Panel A plots the estimated remittance coefficients against overall averaged financial development of the recipient country. The number of countries included in the exercise is 78. Two countries—i.e., Mauritania and West Bank and Gaza — are excluded because of the lack of data availability. Panel B plots the estimated remittance coefficients against institutional quality proxied by regulatory quality. The number of countries included in the exercise is 77. Similarly, the three countries—i.e., Congo, Dem. Rep., Romania , and West Bank and Gaza — excluded in the analysis is due to the lack of data availability. Countries and their corresponding codes are available in Table E.4 in the Appendix.

The financial development index we employ for analysis ranges from 0–1, where zero (one) represents lowest (highest) levels of financial development. The institutional quality variable is

[25]There is a general consensus in the remittance literature that the financial development and institutional quality are major players in mediating the impact of remittance on output, particularly, in the long run (e.g., Abdih et al., 2012; Catrinescu et al., 2009; Giuliano and Ruiz-Arranz, 2009; Mundaca, 2009; Sobiech, 2019).

proxied by regulatory quality and it measures the perceptions of the ability of the government to formulate and implement sound policies and regulations that permit and promote private sector development. The measure ranges between -2.5– 2.5 where higher values correspond to better outcomes and higher levels of institutional quality. For the exercise that follow, we focus primarily on estimates from the DOLS estimation that accounts for cross sectional dependence in Table 3.

Figure 2 depicts the dispersion of the individual remittance coefficient and how it relates to financial development (Panel A) and institutional quality (Panel B). In panel A we observe a weak inverted U-shaped curve suggesting a mild nonlinear relationship between overall financial development and the estimated remittance-output. Specifically, the inverted U-shape reveal that at early stages or low levels, financial development may be positively associated with the impact of remittances on output. This finding is consistent with the notion that remittances and financial development are substitutes at low levels of financial development and confirms the finding that remittances foster growth only at low levels of financial development (Giuliano and Ruiz-Arranz, 2009; Sobiech, 2019). At the higher levels of financial development, a weak negative association between financial development and the estimated remittance-output relationship is observed. The relationship is however statistically non-significant for the higher levels of financial development.

Panel B of Figure 2 portrays the association between remittance-output estimates and institutional quality proxied by regulatory quality. The figure uncovers an increasing relationship between institutional quality and the estimated impact of remittance on output.[26]. Specifically, the better the ability of the government to formulate and implement sound policies and regulations that permit and promote private sector development the stronger the positive impact of remittance on output. At low levels of good institution, however, the observed relationship is not statistically significant, but the relationship becomes stronger and statistically significant at higher levels of institutional quality. This results are compatible with Catrinescu et al. (2009) which shows that good institutions facilitates a positive remittance-growth relationship. Furthermore, the results support the theoretical viewpoint that in the presence of higher quality of institutions, remittances are invested efficiently.

8 Establishing a causal relationship

In addition to uncovering an important long-run economic relationship between remittances and output, the results in the previous section also implies a long-run causality in at least one direction. Nonetheless, it does not show the direction of the long run causality. Assessing long-run association between remittance and real GDP without any insight into the potential causal

[26]It is worth mentioning that this results also holds when we employ government effectiveness and rule of law as measures of institutional quality.

link between these two variables does not inform policy strategies and implications. More precisely, the positive relationship between remittance and output could also suggest that robust aggregate economic activity in recipient countries induces higher remittances. [27]. To examine whether there is a potential causal relationship, we utilize a causality test by Dumitrescu and Hurlin (2012, , henceforth DH). Unlike the standard Granger causality test, the DH- test allows the use of unbalanced data and considers two dimensions of heterogeneity— i.e., the heterogeneity of the regression model used to test the Granger causality, and the heterogeneity of the causal relationship. For a given pair of economic variables, X and Y, the null hypothesis of the DH-test is that X does not homogenously cause Y.

Table 10: Causality Tests

Null Hypothesis	Zbar-Stat
Remittances do not homogenously cause Output	2.109**
Output does not homogenously cause Remittances	2.813**

Notes: ** indicates rejection of the null hypothesis that the variable X does not homogenously cause the economic variable Y. The number of lags in the individual regressions was set to 1 as guided by the Akaike information criteria and Hannan-Quinn information criteria. We employ the cross-sectionally demeaned data, which we tested to be stationary.

Table 10 presents the results from the test. The evidence from the table shows the rejection of the null hypothesis. This suggests that there exists a long-run bi-directional causal relationship between remittance and output such that a rise in remittance increases output and that, in turn, higher output induces an increase in remittance inflows. We have discussed earlier, how and why remittance can induce an increase in output. For a possible explanation of why a rise in the recipient country's GDP increase remittance inflows we can rely on the investment motives of remitters as discussed in De et al. (2019). The investment motive suggests that families send migrants to increase the family's income. One can rationalize remittances as a return on the deployment of human capital. To this end, family members then act as agents managing the funds on behalf of the remitter. Thus, if investment is the motive, robust economic activity in the recipient country would increase remittances. Another competing argument is inheritance. The potential inheritance can serve as an enforcement device to encourage migrants to send higher amounts of remittance in the hope of receiving a favourable share of the bequest (Hoddinott, 1994). Hence, one would expect that higher output in the recipient country will increase the value of the bequest, which will in turn lead to more remittances.[28]

[27]For instance, a number of studies including Freund and Spatafora (2005) and Sayan (2006) report that higher output lead to higher remittance inflows.

[28]It also important to mention that our results run counter to studies that find that remittances increase when economic activity is not robust (e.g., Frankel, 2011).

9 Concluding Remarks

There is evidence that remittances are associated with accelerated poverty reduction, improved access to education and health services, and enhanced financial development (e.g., Acosta et al., 2008; Adams Jr and Page, 2005; Aggarwal et al., 2011; Azizi, 2019; Fromentin, 2017). However, the aggregate impact of remittance on output remains ambiguous. This paper contributes to the growing and ongoing discussion on the aggregate economic effects of remittances by investigating the heterogeneous long-run relationship between remittances and output in developing countries. Motivated by a remittance-output steady state relationship derived from a two sector open-economy model, we estimate a bivariate cointegration regression of output and remittance. Intuitively, the model permits the interpretation of the remittance-output coefficient as a trade-off impact of remittance on steady state investment, consumption, labor supply, and trade balance. Empirically, we relax the assumption of a homogeneous remittance-output association often assumed in existing studies and allow the long-run remittance relationship to vary across individual countries.

We find that there exists a strong long run relationship between remittance and output such that a 1 percent increase in remittance is associated with a 0.07 percent permanent increase in real GDP on average. Importantly, when we account for investment— a primary factor of growth— in the cointegration relation, we still find a strong positive relationship between remittances and real GDP. These results suggest that the long-run positive relationship between remittance and real GDP is mediated by channels other than the investment channel. The coefficient of investment in the long run relation is positive. In addition to the latter, the baseline estimation result is robust to a host of sensitivity analysis along the lines of data treatment, periodisation and choice of estimator. Interestingly, country-by-country estimations results uncover a large heterogeneity across countries. More specifically, we find that the associated response of real GDP following a permanent increase in remittance varies from -0.53 percent in Bosnia and Herzegovina to 0.59 percent in Dominican Republic. This large variation in the estimated remittance-output relationship sheds light into why existing studies have found mixed results. Moreover, we find that cross-country differences in financial development and institutional quality proxied by regulatory quality explain the large variation in the estimated remittance-output relationship. This finding suggests that financial development and institutional quality play a mediating role in impacting direction of the relationship between remittance-output.

Our results imply that the implementation of polices that ease the impediments of migrant remittances can result in a permanent increase of real GDP in the recipient countries. For instance, Freund and Spatafora (2008) find that a 1 percentage point reduction in transaction costs raises recorded remittances by 14–23 percent. Our estimates suggest that such a rise in remittances would be associated with a 0.98 – 1.61 percent increase in output in the developing countries. The finding is particularly important because it suggests that making remittances easier and cheaper to send can induce a sustained and permanent increase in the output in the recipient economies. Furthermore, we find that cross-country differences in institution quality is positively correlated to remittance-output relationship across countries— suggesting that a ”good” institutional environment is essential. Specifically, the improvement of institutional quality along the lines of government's ability to formulate and implement policies that promote private sector development would enhance the long run remittance-output relationship.

We show theoretically the channels through which remittance can impact output, and estimate an aggregate impact of remittance, which captures the trade-off effects across these channels. A natural extension of this paper would be to employ recently developed heterogeneous panel structural vector autoregression model in the manner of Pedroni (2013) to uncover the contribution of the individual channels. This approach will not only shed light on the relevant channels that influence the positive remittance-output association, it will also preserve the heterogeneous feature highlighted in this study.

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APPENDICES

A The Full Model

The model in this paper is described in section 2. Here, we present the details on the derivation of key equations and expressions employed for the key relationships in section 2.

A.1 Deriving the intratemporal allocation consumption bundles

$$C = \left[\varphi^{\frac{1}{\phi}}(C^{T})^{\frac{\phi-1}{\phi}} + (1-\varphi)^{\frac{1}{\phi}}(C^{N})^{\frac{\phi-1}{\phi}}\right]^{\frac{\phi}{\phi-1}} \tag{A.1}$$

subject to the budget constraint,

$$C^{T} + PC^{N} = P_{C}C \tag{A.2}$$

Combining first order condition from the optimization problem arising from (A.1) nad (A.2) yields the relation,

$$C^{T} = \frac{\varphi}{1-\varphi}P^{\phi}C^{N} \tag{A.3}$$

substituting Eq. (A.3) into Eq. (A.1) and solving C^{N}, we obtain

$$C^{N} = (1-\varphi)\left(\frac{P}{P_{C}}\right)^{-\phi}C_{t} \tag{A.4}$$

inserting Eq. (A.4) into (**??**) yields

$$C^{T} = \varphi(P_{C})^{\phi}C_{t} \tag{A.5}$$

Furthermore, we can rewrite Eq. (A.4) as $C^{N} = P_{N}'C$. To acheive this, recall that $P_{C} = \left[\varphi + (1-\varphi)P_{N}^{1-\phi}\right]^{\frac{1}{1-\phi}}$. Consequently, with some manipulation we can write $\frac{P_{N}}{P_{C}} = \frac{P_{N}^{1-\phi}}{\varphi+(1-\varphi)P^{1-\phi}}\frac{P_{C}}{P}$. Now, inserting $\frac{P_{N}}{P_{C}}$ into Eq. (A.4) and setting $\alpha_{C} = \frac{(1-\varphi)P^{1-\phi}}{\varphi+(1-\varphi)P_{N}^{1-\phi}}$ and utilising Eq. (A.2) ,we arrive at the following expressions for C^{N} and C^{T}

$$C^{N} = \alpha_{C}\frac{P_{C}}{P}C \tag{A.6}$$

$$C^{T} = (1-\alpha_{C})P_{C}C \tag{A.7}$$

Now using Sherpherd's Lemma, we obtain

$$C^{N} = P_{C}'C \tag{A.8}$$

B Short-run static solution

Following Cardi and Restout (2015), we derive the short-run static solutions.

B.1 Short-run static solutions for households From the first order conditions in Eq. (2.5) and Eq. (2.6), we can solve for short-run static solutions for real aggregation consumption and lobor supply, which are both given as $C = C(\bar{\lambda}, P)$ and $L = L(\bar{\lambda}, P)$ respectively. Here, it straightforward to show that

$$C_{\bar{\lambda}} = -\sigma_C \frac{C}{\lambda} < 0 \tag{B.1}$$

Also, from Eqs. (A.6) and (A.8) $\alpha_C = \frac{P_C' P}{P_C}$. Thus,

$$C_P = -\alpha_C \sigma_C \frac{C}{P} < 0 \tag{B.2}$$

Relatedly, we can derive L_λ and L_P as follows:

$$L_{\bar{\lambda}} = \frac{\sigma_L}{\lambda} L > 0 \tag{B.3}$$

and

$$L_P = \sigma_L L \frac{W_P}{W} \tag{B.4}$$

However, we know from the FOC of the firm, we know that $W(P) = f - k^T f_k$. We therefore have that $W_P = -\frac{k^T h}{k^N - k^T}$. Consequently,

$$L_P = -\sigma_L \frac{L}{W} \frac{k^T h}{k^N - k^T} \lesseqgtr 0 \tag{B.5}$$

Now, substituting the short-run solution for consumption C into the intratemporal allocations between non-tradables and tradable goods, we can solve for C^N and C^T as

$$C^N = C^T(\bar{\lambda}, P), \qquad C^T = C^T(\bar{\lambda}, P) \tag{B.6}$$

We can use Eq. (A.6) and Eq. (A.7) to solve for the partial derivative C^N and C^T with respect to $\bar{\lambda}$ and P. Consequently, it is straightforward to show that

$$C^T_{\bar{\lambda}} = -\sigma_C \frac{C^T}{\bar{\lambda}} < 0 \tag{B.7}$$

$$C^N_{\bar{\lambda}} = -\sigma_C \frac{C^N}{\bar{\lambda}} < 0 \tag{B.8}$$

The derivation of C_P^T and C_P^N involves a few more steps. Recall that $C^N = P_C' C$. Hence, $C_P^N = P_C'' \cdot C + P_C' \cdot C_P$. Substituting Eq. (B.2) into the latter we obtain

$$C_P^N = -\frac{C^N}{P}(\alpha_C \sigma_C - \frac{P_C^{\text{‘}\prime} P}{P_C'})$$

However, from $P_C = \left[\varphi + (1-\varphi)P_N^{1-\phi}\right]^{\frac{1}{1-\phi}}$ we can show that $\frac{P_C^{\text{‘}\prime} P}{P_C'} = \phi(1-\alpha_C)$ so that

$$C_P^N = -\frac{C^N}{P}(\alpha_C \sigma_C + \phi(1-\alpha_C)) < 0 \tag{B.9}$$

From Eq. (A.3) we get $C_P^T = -\alpha_C \sigma_C \frac{C^T}{P} + \phi C^T \frac{P_C' P}{P_C}$. But, from Eqs. (A.6) and Eq. (A.8), $\frac{P_C^{\text{‘}} P}{P_C} = \frac{\alpha_C}{P}$ so that

$$C_P^T = \alpha_C \frac{C^T}{P}(\phi - \sigma_C) \lesseqgtr 0 \tag{B.10}$$

B.2 Short-run static solutions for firms From the firm optimality conditions in Eq. (2.9) and Eq. (2.10), we can solve for $\frac{\partial k_T}{\partial P}$ and $\frac{\partial k_N}{\partial P}$ by taking the derivative with respect to P

$$\theta^T(\theta^T - 1)(k^T)^{\theta^T - 1 - 1}\frac{\partial k^T}{\partial P} = \theta^N (k^N)^{\theta^N - 1} + P\theta^N(\theta^N - 1)(k^N)^{\theta^N - 1 - 1}\frac{\partial k^N}{\partial P} \tag{B.11}$$

$$\theta^T(\theta^T - 1)(k^T)^{\theta^T - 1}\frac{\partial k^T}{\partial P} = (1-\theta^N)(k^N)^{\theta^N} + P\theta^N(1-\theta^N)(k^N)^{\theta^N - 1}\frac{\partial k^N}{\partial P} \tag{B.12}$$

Notice that $f_{kk} = \theta^T(\theta^T - 1)(k^T)^{\theta^T - 1 - 1}$ and $h_{kk} = \theta^N(\theta^N - 1)(k^N)^{\theta^N - 1 - 1}$. Hence, from Eq. (B.11) and Eq. (B.12) we obtain respectively,

$$k^N f_{kk}\frac{\partial k^T}{\partial P} = \theta^N (k^N)^{\theta^N} + P\theta^N(\theta^N - 1)(k^N)^{\theta - 1}\frac{\partial k^N}{\partial P} \tag{B.13}$$

$$-k^T f_{kk}\frac{\partial k^T}{\partial P} = (1-\theta^N)(k^N)^{\theta^N} - P\theta^N(\theta^N - 1)(k^N)^{\theta - 1}\frac{\partial k^N}{\partial P} \tag{B.14}$$

Substracting Eqs. (B.13) and (B.14) and setting $h = (k^N)^{\theta^N}$, we obtain the $\frac{\partial k^T}{\partial P}$

$$\frac{\partial k^T}{\partial P} = \frac{h}{(k^N - k^T) f_{kk}} \tag{B.15}$$

Multiplying Eq. (B.11) through by $P \cdot (k^T)$ and Eq. (B.12) through by P and combining the resulting equations, we obtain

$$\frac{\partial k^N}{\partial P} = \frac{f}{(k^N - k^T)P^2 h_{kk}} \tag{B.16}$$

As previously shown the wage rate is given as

$$W_P = -\frac{k^T h}{k^N - k^T} \tag{B.17}$$

Labor

Recall that $L = L(\bar{\lambda}, P)$, $k^T(P)$ and $k^N(P)$. Thus, utilizing the the resource constraints for labor and capital in Eq. (3.12) and Eq. (3.18), we can solve for traded and non-traded labor such that $L^T = L^T(K, P, \bar{\lambda})$ and $L^N = L^N(K, P, \bar{\lambda})$ where,

$$L^T = \frac{K - Lk^N}{k^T - k^N} \tag{B.18}$$

$$L^N = \frac{Lk^T - K}{k^T - k^N} \tag{B.19}$$

Consequently, we have

$$L^T_{\bar{\lambda}} = \frac{\sigma_C}{\bar{\lambda}} \frac{Lk^N}{k^N - k^T} \lesseqgtr 0 \tag{B.20}$$

$$L^N_{\bar{\lambda}} = \frac{\sigma_C}{\bar{\lambda}} \frac{Lk^T}{k^T - k^N} \lesseqgtr 0 \tag{B.21}$$

$$L^T_K = \frac{1}{k^T - k^N} \lesseqgtr 0 \tag{B.22}$$

$$L^N_K = \frac{1}{k^N - k^T} \lesseqgtr 0 \tag{B.23}$$

$$L^T_P = \frac{1}{(k^N - k^T)^2} \left\{ \frac{L^N f}{P^2 h_{kk}} + \frac{L^T h}{f_{kk}} - \frac{\sigma_L L k^T k^N h}{W} \right\} < 0 \tag{B.24}$$

$$L^N_P = -\frac{1}{(k^N - k^T)^2} \left\{ \frac{L^N f}{P^2 h_{kk}} + \frac{L^T h}{f_{kk}} - \frac{\sigma_L L h (k^T)^2}{W} \right\} > 0 \tag{B.25}$$

where $h_{kk}, f_{kk} < 0$.

Output

Substituting the short-run static solution for capital-labor ratios and labor into the sectoral production functions, we obtain traded and non-traded output as $Y^T = L^T f(k^T)$ and $Y^N = L^T h(k^N)$, respectively as $Y^T = L^T(K, P, \bar{\lambda})$ and $Y^N = L^N(K, P, \bar{\lambda})$. Subsequently, we can derive the following:

$$Y^T_{\bar{\lambda}} = \frac{\sigma_L L}{\lambda}\frac{k^N f}{k^T - k^N} \lesseqgtr 0 \tag{B.26}$$

$$Y^N_{\bar{\lambda}} = -\frac{\sigma_L L}{\lambda}\frac{k^T h}{k^N - k^T} \lesseqgtr 0 \tag{B.27}$$

Consequently, we have

$$Y^T_K = \frac{f}{k^T - k^N} \lesseqgtr 0 \tag{B.28}$$

$$Y^N_K = \frac{h}{k^N - k^T} \lesseqgtr 0 \tag{B.29}$$

$Y^T_P = L^T_P f + L^T f_k k^T_P$ and $Y^T_P = L^N_P h + L^N h_k k^N_P$. Substituting the necessary equations and simplifying yields

$$Y^T_P = \frac{1}{(k^N - k^T)^2}\left\{\frac{L^N f^2}{P^2 h_{kk}} + \frac{L^T P h^2}{f_{kk}} - \frac{\sigma_L L k^T k^N h}{W}\right\} < 0 \tag{B.30}$$

$$Y^N_P = -\frac{1}{(k^N - k^T)^2}\left\{\frac{L^N f^2}{P^3 h_{kk}} + \frac{L^T h^2}{f_{kk}} - \frac{\sigma_L L h^2 (k^T)^2}{W}\right\} > 0 \tag{B.31}$$

B.3 Equilibrium dynamic and solution To formally derive the system in Eq. (3.14), substitute in the necessary static solutions related C^N and Y^N and k^N into Eq. (3.13) and (3.8), respectively to obtain,

$$\dot{P} = [r^* + \delta - h_k(k^N(P))]P \tag{B.32}$$

where from Eq. (3.9) we have $R^k/P = h_k(k^N(P))$

$$\dot{K} = Y^N(K, P, \bar{\lambda}) - C^N(P, \bar{\lambda}) - \delta K \tag{B.33}$$

Linearizing Eqs. (B.34) and (B.33) and simplifying yields[29],

$$\dot{P} = \frac{Y^T_K}{P}(P(t) - \bar{P}) \tag{B.34}$$

$$\dot{K} = (Y^N_K - \delta)(K(t) - \bar{K}) + (Y^N_P - C^N_P)(P(t) - \bar{P}) \tag{B.35}$$

[29] The steady state value of an economic variable X is given by $\bar{X}$

where $-h_{kk}(k^N(P))Pk_p^N = \frac{Y_K^T}{P}$. In matrix form, we obtain

$$\begin{pmatrix} \dot{K} \\ \dot{P} \end{pmatrix} = \underbrace{\begin{pmatrix} (Y_K^N - \delta) & (Y_P^N - C_P^N) \\ 0 & \frac{Y_K^T}{P} \end{pmatrix}}_{\equiv A} \begin{pmatrix} K - \bar{K} \\ P - \bar{P} \end{pmatrix} \tag{B.36}$$

The system above corresponds to the system in Eq. (3.14) and $a_{11} = Y_K^N - \delta$; $a_{12} = Y_P^N - C_P^N$ and $a_{22} = \frac{Y_K^T}{P}$

Notice that with the first term set to the matrix A in Eq. (B.36) and $z = (K, P)'$, we can write the system compactly as $\dot{z} = Az$, which a first order ODE to be solved. The general solution to the system in Eq.(B.36) is given by

$$K - \bar{K} = \omega_1^1 B_1 e^{\nu_1 t} + \omega_1^2 B_2 e^{\nu_2 t} \tag{B.37}$$

$$P - \bar{P} = \omega_2^1 B_1 e^{\nu_1 t} + \omega_2^1 B_2 e^{\nu_2 t} \tag{B.38}$$

where ω_j^i's are eigenvectors. we normalize ω_1^i to 1 as in Cardi and Restout (2015). We now have to formally solve for the eigenvalues ν_i's in Eqs. (B.37) and (B.38). Consider the characteristic equation associated with the matrix A[30]

$$\lambda^2 - \text{Tr}(A)\lambda + \det(A) = 0, \tag{B.39}$$

where the trace of matrix A, $\text{Tr}(A)$ is given by

$$\text{Tr}(A) = Y_K^N + \frac{Y_K^T}{P} - \delta = r^* > 0 \tag{B.40}$$

Specifically, we know that $Y_K^N + \frac{Y_K^T}{P} = \frac{f - Ph}{k^T - k^N}$ and from the FOC of the firm we can show that $\frac{f-Ph}{k^T-k^N} = Ph_{kk} = R^k$. It therefore follows from the steady state version on Eq. (3.8) that $h_k - \delta = r^*$. Also,

$$\det(A) = (Y_K^N - \delta)\frac{Y_K^T}{P} < 0 \tag{B.41}$$

The roots associated with the characteristic with Eq. (B.39) is

$$\nu_i = \frac{1}{2}\left\{\text{Tr}(A) \pm \sqrt{(\text{Tr}(A))^2 - 4 \cdot \det(A)}\right\} \tag{B.42}$$

Using Eqs. (B.40) and (B.41) and simplifying Eq. (B.41) yields

[30] i.e., $\det(A - \lambda I) = 0$, where I is the identity matrix and the scalar λ is an eigenvalue of A.

$$\nu_i = \frac{1}{2}\left\{r^* \pm \left(\frac{Y_K^T}{P} - (Y_K^N - \delta)\right)\right\} \lessgtr 0 \tag{B.43}$$

Assuming the stable eigenvalue to be $\nu_1 < 0$, then the associated eigenvector $\omega^1 = (\omega_1^1, \omega_2^1)$ associated with this eigenvalue is

$$\omega_1^1 = 1, \qquad \omega_2^1 = \frac{\nu_1 - a_{11}}{a_{12}} \tag{B.44}$$

Solution for the stock of foreign assets
Consider the current accoount (CA) balance

$$\dot{B}(t) = r^* B(t) + Y^T(K, P, \bar{\lambda}) - C^T(P, \bar{\lambda}) + Rem \tag{B.45}$$

Linearizing Eq. (B.45) around the steady state yields

$$\dot{B}(t) = r^*(B(t) - \bar{B}) + Y_K^T(K(t) - \bar{K}) + (Y_P^T - C_P^T)(P(t) - \bar{P}) \tag{B.46}$$

Substituting Eqs.(B.37) and (B.38) into Eq. (B.46) and multiply the resulting expression by $\mathrm{e}^{-r^* t}$, we obtain

$$[\dot{B}(t) - r^*(B(t) - \bar{B})]\mathrm{e}^{-r^* t} = Y_K^T \sum_{i=1}^{2} B_i \mathrm{e}^{(\nu_i - r^*)t} + (Y_P^T - C_P^T)\sum_{i=1}^{2} B_i \omega_2^i \mathrm{e}^{(\nu_i - r^*)t} \tag{B.47}$$

Solving the differential equation in Eq.(B.47) yields

$$B - \bar{B} = [(B_0 - \bar{B}) - \psi_1 B_1 - \psi_2 B_2]e^{r^* t} + \psi_1 B_1 \mathrm{e}^{\nu_1 t} + \psi_2 B_2 \mathrm{e}^{\nu_2 t} \tag{B.48}$$

where $\phi_1 = \frac{Y_K^T + (Y_P^T - C_P^T)\omega_2^1}{\nu_1 - r^*}$ and $\phi_2 = \frac{Y_K^T + (Y_P^T - C_P^T)\omega_2^2}{\nu_2 - r^*}$. Notice that in order for the economy to satisfy the intertemporal solvency condition $\lim_{t\to\infty} B\mathrm{e}^{-r^* t} = 0$, then $[(B_0 - \bar{B}) - \psi_1 B_1 - \psi_2 B_2] = 0$ and $B_2 = 0$, so that the linearized version of the economy's intertemporal budget constraint is given as

$$B_0 - \bar{B} = \psi_1(K_0 - \bar{K}), \tag{B.49}$$

where $B(0) = B_0$ and $K(0) = K_0$. Consequently, the stable solution for net foreign assets finally reduces to

$$B(t) - \bar{B} = \psi_1(K(t) - \bar{K}). \tag{B.50}$$

Following Brock and Turnovsky (1994); Cardi and Restout (2015) and empirical evidence suggesting a negative relationship between current account and investment (Glick and Rogoff, 1995), we impose the assumption that $\psi_1 < 0$.[31]

[31] See (Cardi and Restout, 2015) for a detailed discussion.

B.4 Long-run analysis

The steady state system can be summarized as follows,

$$h_{kk}(k^N(\bar{P})) = r^* + \delta \tag{B.51}$$

$$Y^N(\bar{K},\bar{P},\bar{\lambda}) - C^N(\bar{P},\bar{\lambda}) - \delta\bar{K} = 0 \tag{B.52}$$

$$r^*\bar{B} + Y^T(\bar{K},\bar{P},\bar{\lambda}) - C^T(\bar{P},\bar{\lambda}) = -\overline{Rem} \tag{B.53}$$

$$B_0 - \bar{B} = \psi_1(K_0 - K) \tag{B.54}$$

We can further substitute Eq. (B.54) into Eq. (B.53) and totally differentiate the resulting system to obtain

$$\begin{pmatrix} h_{kk}k^N_P & 0 & 0 \\ (Y^N_P - C^N_P) & (Y^N_K - \delta) & (Y^N_{\bar{\lambda}} - C^N_{\bar{\lambda}}) \\ (Y^T_P - C^T_P) & (Y^T_K + r^*\psi_1) & (Y^T_{\bar{\lambda}} - C^T_{\bar{\lambda}}) \end{pmatrix} \begin{pmatrix} \mathrm{d}\bar{P} \\ \mathrm{d}\bar{K} \\ \mathrm{d}\bar{\lambda} \end{pmatrix} = \begin{pmatrix} 0 \\ 0 \\ -\mathrm{d}\bar{Rem} \end{pmatrix} \tag{B.55}$$

The determinant, which we denote as $\Delta > 0$, of the matrix of coefficient is given by:

$$\Delta = h_{kk}k^N_P[(Y^N_K - \delta)(Y^T_{\bar{\lambda}} - C^T_{\bar{\lambda}}) - (Y^T_K + r^*\psi_1)(Y^T_{\bar{\lambda}} - C^T_{\bar{\lambda}}] \tag{B.56}$$

Solving the system in B.55 yields the following

$$\frac{\mathrm{d}\bar{P}}{\mathrm{d}Rem} = 0, \tag{B.57}$$

$$\frac{\mathrm{d}\bar{K}}{\mathrm{d}Rem} = \frac{h_{kk}k^N_p}{\Delta}(Y^N_\lambda - C^N_\lambda) \lesseqgtr 0, \tag{B.58}$$

$$\frac{\mathrm{d}\bar{\lambda}}{\mathrm{d}Rem} = -\frac{h_{kk}k^N_p}{\Delta}(Y^N_K - \delta) < 0, \tag{B.59}$$

C Countries employed in the empircal study and number of observations

Table C.1: Countries included in the panel

Country	$t = 1$	$t = T$	Observations	Country	$t = 1$	$t = T$	Observations
Albania	1992	2014	23	Lesotho	1975	2014	40
Algeria	1970	2014	45	Macedonia, FYR	1996	2014	19
Armenia	1995	2014	20	Madagascar	1974	2014	41
Bangladesh	1976	2014	39	Malawi	1994	2014	21
Belarus	1993	2014	22	Mali	1975	2014	40
Belize	1984	2014	31	Mauritania	1975	1998	24
Benin	1974	2014	41	Mauritius	1994	2014	21
Bolivia	1976	2014	39	Mexico	1979	2014	36
Bosnia and Herze	1998	2014	17	Moldova	1995	2014	20
Botswana	1975	2014	40	Mongolia	1998	2014	17
Brazil	1975	2014	40	Morocco	1975	2014	40
Bulgaria	1996	2014	19	Mozambique	1980	2014	35
Cabo Verde	1980	2014	35	Myanmar	2000	2014	15
Cambodia	1993	2014	22	Namibia	1990	2014	25
Cameroon	1979	2014	36	Nepal	1993	2014	22
China	1982	2014	33	Nicaragua	1992	2014	23
Colombia	1970	2014	45	Niger	1974	2014	41
Congo, Dem. Rep.	2000	2014	15	Nigeria	1977	2014	38
Costa Rica	1977	2014	38	Pakistan	1976	2014	39
Dominica	1977	2014	38	Paraguay	1975	2014	40
Dominican Republ	1970	2014	45	Peru	1990	2014	25
Egypt, Arab Rep.	1977	2014	38	Philippines	1977	2014	38
El Salvador	1976	2014	39	Romania	1994	2014	21
Eswatini	1974	2014	41	Russian Federati	1994	2014	21
Ethiopia	1981	2014	34	Sao Tome and Pri	2001	2014	14
Fiji	1979	2014	36	Senegal	1974	2014	41
Georgia	1997	2014	18	South Africa	1970	2014	45
Ghana	1979	2014	36	Sri Lanka	1975	2014	40
Grenada	1986	2014	29	St. Vincent and	1986	2014	29
Guatemala	1977	2014	38	Sudan	1977	2014	38
Haiti	1998	2014	17	Tanzania	1995	2014	20
Honduras	1974	2014	41	Thailand	1975	2014	40
India	1975	2014	40	Togo	1974	2014	41
Indonesia	1983	2014	32	Tunisia	1976	2014	39
Iran, Islamic Re	1993	2014	22	Turkey	1974	2014	41
Jamaica	1976	2014	39	Uganda	1999	2014	16
Jordan	1975	2014	40	Ukraine	1996	2014	19
Kazakhstan	1995	2014	20	Vietnam	2000	2014	15
Kyrgyz Republic	1993	2014	22	West Bank and Ga	1995	2014	20
Lao PDR	1984	2014	31	Yemen, Rep.	1990	2014	25

D Summary Statistics

Table D.2: Summary Statistics by Country, GDP (in natural log)

Country	Mean	Min	Max	Std. Dev.	Country	Mean	Min	Max	Std. Dev.
Albania	22.83	22.15	23.27	0.36	Lesotho	20.98	19.94	21.77	0.49
Algeria	25.25	24.27	25.93	0.42	Macedonia, FYR	22.79	22.54	23.05	0.17
Armenia	22.60	21.93	23.13	0.43	Madagascar	22.55	22.29	22.99	0.23
Bangladesh	24.74	23.93	25.71	0.53	Malawi	22.35	21.85	22.84	0.28
Belarus	24.33	23.75	24.87	0.39	Mali	22.39	21.65	23.21	0.48
Belize	20.53	19.62	21.16	0.49	Mauritania	21.40	21.15	21.66	0.15
Benin	22.04	21.32	22.87	0.47	Mauritius	22.75	22.29	23.17	0.28
Bolivia	23.22	22.82	23.92	0.33	Mexico	27.38	26.92	27.80	0.27
Bosnia and Herze	23.40	23.00	23.60	0.20	Moldova	22.28	21.96	22.68	0.23
Botswana	22.42	20.79	23.52	0.79	Mongolia	22.49	22.03	23.16	0.37
Brazil	27.94	27.32	28.52	0.33	Morocco	24.61	23.76	25.41	0.48
Bulgaria	24.45	24.15	24.69	0.21	Mozambique	22.13	21.26	23.32	0.67
Cabo Verde	20.27	19.08	21.30	0.78	Myanmar	24.27	23.49	24.91	0.46
Cambodia	22.68	21.92	23.42	0.48	Namibia	22.81	22.33	23.36	0.31
Cameroon	23.63	23.16	24.18	0.27	Nepal	23.23	22.76	23.67	0.27
China	28.26	26.69	29.75	0.92	Nicaragua	22.68	22.27	23.11	0.25
Colombia	25.74	24.83	26.58	0.48	Niger	22.01	21.60	22.73	0.30
Congo, Dem. Rep.	23.65	23.32	24.10	0.25	Nigeria	25.89	25.34	26.84	0.45
Costa Rica	23.73	23.08	24.49	0.46	Pakistan	25.25	24.24	26.05	0.53
Dominica	19.63	18.94	20.04	0.31	Paraguay	23.20	22.23	23.93	0.43
Dominican Republ	23.88	22.76	24.89	0.58	Peru	25.31	24.79	25.92	0.35
Egypt, Arab Rep.	25.42	24.48	26.20	0.51	Philippines	25.48	24.95	26.25	0.37
El Salvador	23.37	23.02	23.75	0.23	Romania	25.65	25.37	25.94	0.21
Eswatini	21.51	20.41	22.36	0.62	Russian Federati	27.81	27.42	28.17	0.27
Ethiopia	23.36	22.77	24.51	0.51	Sao Tome and Pri	18.95	18.62	19.29	0.22
Fiji	21.62	21.30	22.01	0.22	Senegal	22.75	22.20	23.42	0.36
Georgia	22.94	22.49	23.39	0.31	South Africa	26.19	25.65	26.75	0.31
Ghana	23.54	22.84	24.52	0.49	Sri Lanka	24.00	23.08	25.01	0.56
Grenada	20.21	19.75	20.55	0.26	St. Vincent and	20.06	19.56	20.39	0.27
Guatemala	23.97	23.49	24.59	0.35	Sudan	24.11	23.45	24.96	0.52
Haiti	22.63	22.56	22.77	0.06	Tanzania	23.84	23.32	24.43	0.36
Honduras	22.92	22.15	23.62	0.42	Thailand	25.79	24.54	26.67	0.66
India	27.15	26.17	28.39	0.68	Togo	21.52	21.09	22.07	0.27
Indonesia	26.83	26.06	27.57	0.43	Tunisia	23.86	23.05	24.59	0.47
Iran, Islamic Re	26.63	26.31	26.94	0.22	Turkey	26.75	25.93	27.66	0.49
Jamaica	23.12	22.80	23.37	0.20	Uganda	23.47	22.99	23.94	0.33
Jordan	23.22	21.96	24.11	0.56	Ukraine	25.49	25.16	25.76	0.22
Kazakhstan	25.34	24.80	25.94	0.41	Vietnam	25.28	24.84	25.70	0.28
Kyrgyz Republic	22.04	21.62	22.49	0.26	West Bank and Ga	22.67	22.29	23.11	0.27
Lao PDR	21.97	21.14	22.99	0.57	Yemen, Rep.	23.76	23.19	24.15	0.30

Table D.3: Summary Statistics by Country, Remittances (in natural log)

Country	Mean	Min	Max	Std. Dev.	**Country**	Mean	Min	Max	Std. Dev.
Albania	20.93	20.36	21.22	0.25	Lesotho	20.26	19.74	20.60	0.22
Algeria	21.02	19.99	22.08	0.59	Macedonia, FYR	19.29	18.36	19.80	0.49
Armenia	20.32	18.82	21.47	1.04	Madagascar	16.41	11.20	19.76	2.32
Bangladesh	21.37	17.64	23.35	1.30	Malawi	15.86	13.84	17.78	1.49
Belarus	19.29	13.33	20.72	2.03	Mali	19.07	17.81	20.46	0.67
Belize	17.41	16.57	18.19	0.52	Mauritania	15.60	13.74	18.11	1.06
Benin	18.69	17.05	19.41	0.46	Mauritius	19.27	18.88	19.59	0.18
Bolivia	17.75	14.56	21.05	2.21	Mexico	23.00	20.28	24.05	0.81
Bosnia and Herze	21.68	21.32	22.30	0.29	Moldova	20.35	14.66	21.34	1.55
Botswana	18.17	16.67	19.12	0.64	Mongolia	18.93	16.71	20.02	0.97
Brazil	21.10	18.56	22.94	1.41	Morocco	21.82	20.94	22.76	0.58
Bulgaria	20.64	18.39	21.84	1.25	Mozambique	17.85	16.59	18.75	0.51
Cabo Verde	18.33	17.37	19.05	0.55	Myanmar	19.30	18.02	21.35	0.96
Cambodia	18.45	16.32	19.63	1.13	Namibia	17.16	16.65	18.34	0.59
Cameroon	17.76	15.76	19.37	0.96	Nepal	20.50	18.30	22.45	1.57
China	21.92	18.98	24.72	2.02	Nicaragua	19.75	17.09	20.77	1.06
Colombia	20.94	18.76	22.64	1.27	Niger	17.20	15.82	19.10	0.93
Congo, Dem. Rep.	16.36	13.75	18.44	1.02	Nigeria	20.72	16.23	24.34	2.90
Costa Rica	18.32	15.74	20.49	1.67	Pakistan	22.15	20.77	23.40	0.60
Dominica	16.95	16.01	17.70	0.30	Paraguay	18.95	16.79	20.09	1.07
Dominican Republ	20.68	18.12	22.27	1.24	Peru	20.93	19.08	21.73	0.75
Egypt, Arab Rep.	22.71	21.72	23.48	0.42	Philippines	22.60	20.89	23.95	1.02
El Salvador	20.86	17.85	22.09	1.11	Romania	19.78	17.10	21.91	1.41
Eswatini	18.40	16.09	19.54	0.70	Russian Federati	22.40	22.06	23.08	0.21
Ethiopia	17.72	15.27	21.08	1.67	Sao Tome and Pri	14.99	13.76	16.79	0.94
Fiji	17.84	15.71	19.14	1.00	Senegal	19.47	16.93	21.35	1.11
Georgia	20.50	19.89	21.27	0.48	South Africa	19.46	18.23	20.80	0.91
Ghana	17.55	13.84	21.56	2.13	Sri Lanka	20.96	16.99	22.59	1.31
Grenada	17.42	16.36	18.27	0.44	St. Vincent and	17.00	14.91	17.57	0.56
Guatemala	19.63	12.44	22.29	2.60	Sudan	20.17	18.19	21.78	1.05
Haiti	20.98	20.14	21.30	0.32	Tanzania	17.42	14.58	19.83	1.54
Honduras	18.70	14.86	21.89	2.73	Thailand	21.14	17.84	22.54	1.07
India	23.09	20.75	25.04	1.26	Togo	17.80	16.17	19.72	1.25
Indonesia	21.25	17.06	23.11	1.56	Tunisia	20.66	19.59	21.58	0.55
Iran, Islamic Re	21.30	20.74	22.58	0.50	Turkey	22.27	20.90	23.12	0.66
Jamaica	20.51	19.10	21.57	0.86	Uganda	20.31	19.74	20.63	0.30
Jordan	21.52	19.86	22.38	0.57	Ukraine	20.63	16.30	22.75	2.32
Kazakhstan	19.26	18.59	20.35	0.53	Vietnam	22.42	21.51	22.96	0.46
Kyrgyz Republic	18.53	14.25	21.29	2.52	West Bank and Ga	20.78	20.12	21.31	0.35
Lao PDR	16.25	14.13	18.63	1.47	Yemen, Rep.	21.57	20.87	22.11	0.33

E Country codes

Table E.4: Country and corresponding country code

country	code	country	code
Albania	ALB	Lesotho	LSO
Algeria	DZA	Macedonia, FYR	MKD
Armenia	ARM	Madagascar	MDG
Bangladesh	BGD	Malawi	MWI
Belarus	BLR	Mali	MLI
Belize	BLZ	Mauritania	MRT
Benin	BEN	Mauritius	MUS
Bolivia	BOL	Mexico	MEX
Bosnia and Herzegovina	BIH	Moldova	MDA
Botswana	BWA	Mongolia	MNG
Brazil	BRA	Morocco	MAR
Bulgaria	BGR	Mozambique	MOZ
Cabo Verde	CPV	Myanmar	MMR
Cambodia	KHM	Namibia	NAM
Cameroon	CMR	Nepal	NPL
China	CHN	Nicaragua	NIC
Colombia	COL	Niger	NER
Congo, Dem. Rep.	COD	Nigeria	NGA
Costa Rica	CRI	Pakistan	PAK
Dominica	DMA	Paraguay	PRY
Dominican Republic	DOM	Peru	PER
Egypt, Arab Rep.	EGY	Philippines	PHL
El Salvador	SLV	Romania	ROU
Eswatini	SWZ	Russian Federation	RUS
Ethiopia	ETH	Sao Tome and Principe	STP
Fiji	FJI	Senegal	SEN
Georgia	GEO	South Africa	ZAF
Ghana	GHA	Sri Lanka	LKA
Grenada	GRD	St. Vincent and the Grenadines	VCT
Guatemala	GTM	Sudan	SDN
Haiti	HTI	Tanzania	TZA
Honduras	HND	Thailand	THA
India	IND	Togo	TGO
Indonesia	IDN	Tunisia	TUN
Iran, Islamic Rep.	IRN	Turkey	TUR
Jamaica	JAM	Uganda	UGA
Jordan	JOR	Ukraine	UKR
Kazakhstan	KAZ	Vietnam	VNM
Kyrgyz Republic	KGZ	West Bank and Gaza	PSE
Lao PDR	LAO	Yemen, Rep.	YEM

F Density Estimates

As part of the robustness exercise, we follow Eberhardt and Teal (2010), we investigate the density estimates for the estimated country-specific remittance-output coefficients from our baseline results in Table 3. The density estimates serves as a diagnostic testing and allows us to formally check whether any significant outlier is driving our baseline results, thereby complementing our findings in section 5.2. Additionally, it permits us to observe first hand the distribution of the estimated country-specific remittance-output coefficient, highlighting the variation of the remittance-output relationship across countries.

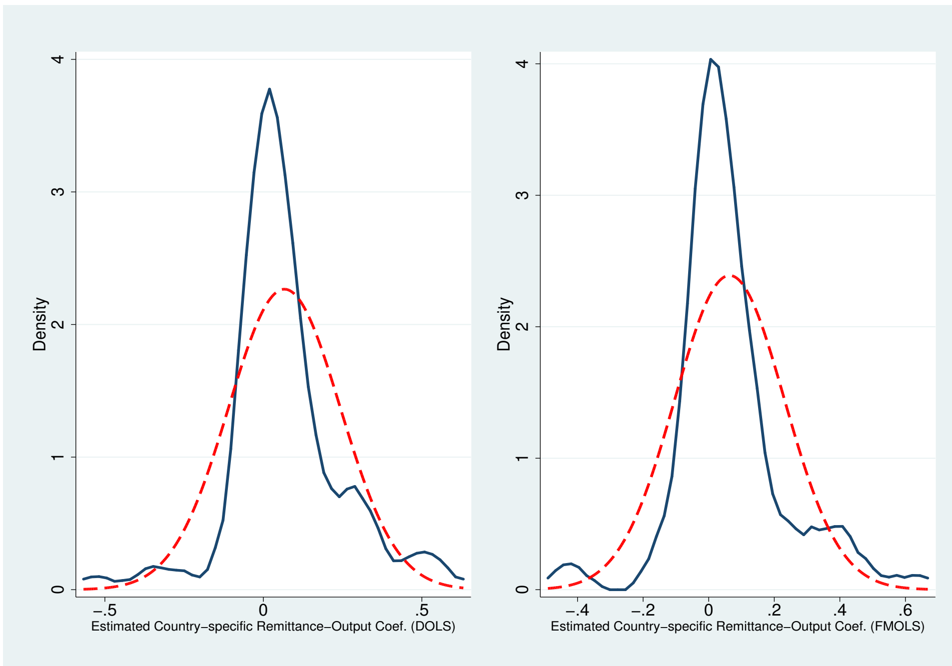

Figure 3: In each case, the kernel is the epanechnikov with automatic bandwidth selection. For the FMOLS the bandwidth is 0.0374 and the DOLS it is 0.0415. The red dashed line is the normal distribution and the blue thick line is the estimated Kernel Density. Recall that the estimated mean for the DOLS was 0.0664 and that of the FMOLS is 0.0638.

Figure 3 reports the kernel density estimates for the country-by-country DOLS and FMOLS estimates in Table 3. Evidently, both panels in the figure show that the distribution of the remittance-output coefficients are symmetric around their means of 0.0664 (DOLS) and 0.0638 (FMOLS), and approximately Gaussian. This further confirms our earlier conclusion that no significant outliers drive our resutls. Moreover, it is clear from the figure that the estimated remittance-output coefficents vary across countries. This implies that an assumption of a homogenous relationship between remittances and output across all countries is a strong one as it may not accurately capture the relevance of a heterogenous relationship.