Exploring the Effects of Remittances on Lithuanian Economic Growth

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We explore the effect of remittances on the output per worker in Lithuania over the sample period 1980–2012. We use the augmented Solow approach and the ARDL bounds procedure to examine the cointegration relationship and subsequently estimate short-run and long-run effects, and the causality nexus. The data available in the World Bank database for the required variable are relatively small for the analysis. To overcome this constraint, we plot data of the respective variables and use appropriate trend functions to approximate data for missing years. For remittances, we use the exponential trend function as the best fit to approximate data for remittances over the period of 1980–1992 and 2012. Similar approach is used to approximate data for GDP (at current prices) for the period of 1980–1989 and gross fixed capital formation for the period of 1980–1994 to build the capital stock data. Further, we use the polynomial trend function to approximate data for GDP (constant 2005 USD) for the period of 1980–1994. Within the necessary caveats, our results show that output per worker, capital per worker and remittances are cointegrated, and remittances contribute about 0.02 % and 0.04 % to output per worker in short-run and long-run, respectively, and the capital share is around 0.50. We note a bidirectional causation and hence a mutually reinforcing effect between capital per worker and output per worker, and a unidirectional causation from remittances to output per worker, duly supporting the remittances-led growth hypothesis in Lithuania. However, our results are not unambiguous due to data constraints. In this regard, further research will supplement and give more insights to the results and outcomes of this paper.

Keywords: remittances, capital per worker, cointegration, the ARDL approach, the Granger causality tests, economic growth, Lithuania.

Introduction

The World Bank classifies Lithuania as a high income country, with a reported GDP per capita of PPP $22,566 in 2013. Between 1990 and 2012, the net out migration was summed up to 425,579 people, which represents just over 14.2 % of the total population.¹ According to Statistics Lithuania, the total number of official and unofficial emigrants aggregated to just 615,000 people between 1990 and 2011 (IOM, 2011). Since independence of Lithuania in 1991, around 30 % of the population between the ages 15-64 years left the country. Furthermore, 55 % of the emigrants were between the age of 20–35 years and 50 % of these migrants had upper secondary education and 25 % had completed higher and professional education (IOM, 2011). Subsequently, it can be concluded, that the average Lithuanian emigrant is highly educated and young. Lithuanian emigrants leave for work mostly to the UK (33 %), Ireland (16 %), the USA (11 %), Germany (8 %), Russian Federation (7 %), Belarus (5 %), Spain (4 %) and Denmark (3 %) (Rakauskiene & Ranceva, 2012).

¹ The authors' own calculations are based on Rakauskiene and Ranceva (2012) and the European migration network (EMN) http://123.emn.lt/en/ general-trends/migration-10-years-overview.

According to Eurostat (2013), Lithuania has realized the most negative population growth rate (-1.06 %) and the second-highest emigration rate (0.71 %) in the European Union in 2012. Furthermore, Rakauskiene & Ranceva (2013) and Eurostat (2013) state that Lithuania has the highest rate of change of the share of population, which is aged 65 years or older (7.1 percentage points between 1991 and 2011). However, it is also worthwhile to note that although the number of immigrants is low (on average, 6000 people per year), eight out of ten of them are nationals and until now mostly single persons emigrated. Moreover, it looks like the share of emigrating families have increased (IOM, 2011) and close to 85 % of the emigrants were unemployed for at least a year before they emigrated. The more important fact is that Lithuania has also experienced growth in remittances inflow (Thaut, 2009). Krupickaite & Poviliunas (2012) highlight that remittances in 2010 accounted for 4.3 % of GDP and 23.9 % of the entire net salary fund in Lithuania.

In one study, Kasnauskiene & Buzyte (2011) examine the impacts of remittances on the economic growth of Lithuania and Poland. Their findings show that in case of Poland, remittances have positive influence on economic growth. However, in case of Lithuania, they find that workers’ remittances, which account for over 80 percent of
total remittances, are not statistically significant in the model and have a negative impact on the GDP growth per capita. On the other hand, Kasnauskiene & Stumbryte (2012) use a dynamic panel approach and find that on average, remittances contribute about 0.0348% to GDP. In another study, Damuliene (2013) find a strong correlation between private remittances to Lithuania and the number of emigrants, and hence anticipates that remittances increase aggregate demand, domestic consumption and GDP growth. In contrast to the critical view on Lithuanian emigration in the literature (c.f. Rakauskiene & Ranceva, 2012, Thaut, 2009, Krupickaitė & Povilaitienė, 2012) and in the media, in this article, and the mixed results, we hypothesize a positive impact of emigration for the Lithuanian economy. More specifically, we focus on the effects of remittances on the Lithuanian economic growth using a different method and approach.

From the remittances-growth nexus literature, we find very few studies that have looked at the role of remittances in Lithuania (Kasnauskiene & Buzyte, 2011; Kasnauskiene & Stumbryte, 2012; Damuliene, 2013). One of the reasons for this gap in the literature is the data limitations. To overcome this hurdle, we make a modest attempt to overcome the data limitations and examine the cointegration, short-run and long-run effects, and causality nexus. In what follows, we use the augmented Solow framework (Solow, 1956) and apply the ARDL bounds approach (Pesaran, et al., 2001) to examine cointegration and short-run and long-run effects. In addition, we utilise the Toda and Yamamoto (1995) approach to examine the causality nexus (Rao, 2007). Briefly, we examine the plausible effects of remittances on the output per worker in Lithuania over the period of 1980–2012.

Notably, the number of studies focussing on the impact of remittances on economic activities is growing. Personal remittances (formerly known as workers’ remittances) received are defined as a composition of personal transfers and compensation of employees. Personal transfers include current transfers in cash or in kind received by resident households from non-resident households. Compensation of employees includes income of border, seasonal and other short-term workers who are employed in a country where they are not resident, and of residents employed by non-resident entities (World Bank, 2013). It is argued that remittances have a welfare enhancing effect, particularly when they support consumption, capital investment, education and human development, entrepreneurship, and poverty reduction efforts (Ratha, 2007; Buch & Kuckulenz, 2010; Rao & Hassan, 2012). Empirically, it has been shown that remittances have both a growth enhancing and a poverty-reducing potential. For instance, Adams and Page (2005) study 71 developing countries with the aim to analyse the effects of migration and remittances on inequality and poverty. Their results show that both international migration and remittances significantly reduce the level, depth and severity of poverty in developing countries. Nevertheless, the remittances-led growth (RLG) hypothesis has shown mixed results. For example, Pradhan et al., (2002) examine the effect of workers’ remittances on economic growth in a sample of 39 developing countries using a panel data from the period of 1980–2004 and a standard growth model. Their results show a positive impact of remittances on growth.

In a study, Chami et al., (2003) consider the role of remittances in development and economic growth by constructing a framework which connects the motivational (altruistic) aspects of remittances to the impacts on economic activities, and find that remittances in fact have a negative effect on economic growth largely due to the moral hazard problem in remittances. Gupta et al., (2009) assess the effect of remittances in Sub-Saharan Africa within the context of financial development and poverty reduction and find that remittances have a direct poverty-mitigating effect and the potential to support financial development. (Giuliano & Ruiz-Arranz, 2009) explore the links between remittances and growth within the context of financial development for 100 countries and find remittances can boost growth in countries with a less developed financial system by providing an alternative way to finance investment and help overcome liquidity constraints. Acosta (2008) studies 10 countries in Latin America and the Caribbean (LAC) and finds remittances support growth and reduce inequality and poverty. Mundaca (2009) analyzes the effect of workers’ remittances and financial intermediation on economic growth with a panel data for selected countries in the LAC and states that remittances when used appropriately with effective financial intermediation could result in growth possibilities. Nyamongo et al., (2012) investigate the role of remittances and financial development on economic growth in a panel of 36 countries in Africa over the period of 1980–2009. They find, inter alia, (a) remittances are an important source of growth for African countries; (b) volatility of remittances has a negative effect on growth; (c) remittances seem to complement financial development; and (d) financial development is a weak contributor to growth. Contrary to Giuliano & Ruiz-Arranz (2009), Bettin & Zazzaro (2012) show that in the countries where financial system functions effectively, the RLG hypothesis is highly plausible relative to the countries where banking systems are weak.

On the contrary, some studies find remittances have a negative effect on growth. For instance, Rao & Takirua (2010) examine the plausible sources of growth in a small island economy of Kiribati using the general-to-specific (GETS) technique and find remittances have a long-run negative effect. The rest of the article is organized as follows. In Section 2, we provide the econometric modelling, estimation techniques and results. In section 3, we conclude our results.

Econometric Modelling and Estimation Techniques

Framework

For the purpose of modeling and analysis, we use an approach introduced by Sturm et al. (1998) and Rao (2010) which is related to the augmented Solow (Solow, 1956) framework.2 The initial equation is defined as:

2 Note that if a Cobb-Douglas function is used, the form of neutrality of technical progress (Harrod-neutral, Solow-neutral, and Hicks-neutral) does not play a role since the effect remains the same.
where \( A = \text{stock of technology and } k = \text{capital per worker, and } \alpha \) is the capital share. The Solow model assumes that the evolution of technology is given by:

\[
\Phi_t = A_0 e^{g_t}
\]

where \( A_0 \) is the initial stock of knowledge and \( t \) is time. Next, we introduce remittances, \( RML_k \) as shift variable:

\[
\Psi_t = f(RML_k)
\]

The effect of \( RML_k \) on total factor productivity (TFP) can be captured when the latter is entered as a shift variable into the production function (c.f. Rao, 2010). Subsequently we identify \( RM_t^\beta \) (where \( \beta \in [0,1] \) represents elasticity of remittances) as part of the stock of technology and redefine \( A_t \) as follows:

\[
A_t = \Phi_t \Psi_t = A_0 e^{g_t} RM_t^\beta
\]

where \( A_0 \) is the initial stock of knowledge, \( g \) refers to the growth of technology over time \( t \), and hence, \( e^{g_t} \) includes other catch-all factors. Hence,

\[
y_t = (A_0 e^{g_t} RM_t^\beta) k_t^\alpha
\]

where \( y_t \) is the output per worker.

### Data

We use a perpetual inventory method to build the data for capital stock. We assume depreciation rate (8) of 0.08 and initial capital stock \( (K_0) \) as 0.50 times of the real GDP of 1980 in constant USD (US dollars at 2005 prices). The gross fixed capital formation in constant 2005 USD is used as a proxy for aggregate investment \( (I_t) \). Hence, \( K_t = (1-\delta)K_{t-1} + I_t \) 

The data on the labour stock are estimated using labor force data and redefined output per worker as \( g_y \) includes 

\[
y_t = \Phi_y \Psi_t = A_0 e^{g_y} (RM_t^\beta)
\]

where \( A_0 \) is the initial stock of knowledge, \( g \) refers to the growth of technology over time \( t \), and hence, \( e^{g_y} \) includes other catch-all factors. Hence,

\[
y_t = (A_0 e^{g_y} (RM_t^\beta)) k_t^\alpha
\]

where \( y_t \) is the output per worker.

### ARDL bounds procedure

Next, we specify the ARDL specifications as (6)-(8) below. Note that each equation has a dummy variable (\( Dum_i \)) associated, which represents the cumulative structural breaks in the series and is identified by applying the Zivot & Andrews (2002) unit root test with single structural break test. Including the \( Dum_i \), therefore provides a relatively more robust computation of bound statistics and also short-run and long-run run results.

\[
\Delta y_t = \beta_{10} + \beta_{11} y_{t-1} + \beta_{12} Lk_{t-1} + \beta_{13} LRM_{t-1} + \alpha_{10} Dum_t + \sum_{j=1}^{p} \alpha_{1j} \Delta y_{t-j} + \sum_{j=1}^{p} \alpha_{1j} \Delta Lk_{t-j} + \sum_{j=1}^{p} \alpha_{1j} \Delta LRM_{t-j} + \epsilon_{1t}
\]

\[
\Delta Lk_t = \beta_{20} + \beta_{21} y_{t-1} + \beta_{22} Lk_{t-1} + \beta_{23} LRM_{t-1} + \alpha_{20} Dum_t + \sum_{j=1}^{p} \alpha_{2j} \Delta y_{t-j} + \sum_{j=1}^{p} \alpha_{2j} \Delta Lk_{t-j} + \sum_{j=1}^{p} \alpha_{2j} \Delta LRM_{t-j} + \epsilon_{2t}
\]

\[
\Delta LRM_t = \beta_{30} + \beta_{31} y_{t-1} + \beta_{32} Lk_{t-1} + \beta_{33} LRM_{t-1} + \alpha_{30} Dum_t + \sum_{j=1}^{p} \alpha_{3j} \Delta y_{t-j} + \sum_{j=1}^{p} \alpha_{3j} \Delta Lk_{t-j} + \sum_{j=1}^{p} \alpha_{3j} \Delta LRM_{t-j} + \epsilon_{3t}
\]

The autoregressive distributed lag (ARDL) approach is used because it is relatively simple and recommended for a small sample size (Ghatak & Siddiki, 2001; Pesaran, et al., 2001). To examine the cointegration based on the computed F-statistics, it is recommended to use the critical bounds from Narayan (2005), which are specifically constructed for a small sample size. The critical bounds of Pesaran et al., (2001), however, are suitable in cases when the sample size exceeds 80. Although, one may not test for unit roots and investigate cointegration thus overlooking the order of integration, we need to conduct the unit root tests for a couple of reasons. First, to ensure that the series are indeed I(0) and/or I(1) in order to apply the ARDL bounds procedure instead other approach such as ordinary least squares (OLS) method which is not recommended for variables in the presence of unit root; second, examining the unit root provides information on the maximum lags which are useful when performing the Toda & Yamamoto, (1995) non-Granger causality procedure. Therefore, we use the augmented Dickey-Fuller (ADF), Phillips-Perron (PP) and Kwiatkowski-Phillips-Schmidt-Shin (KPSS) tests to examine the time series properties of the variables and compute the unit root statistics.
Based on these standard tests, we conclude that all variables are stationary at most in their first differences (Table 2) duly confirming the maximum order of integration is one. Furthermore, we use the Zivot and Andrews (2002) unit root test to determine structural breaks in the series. Importantly, the presence of structural breaks can influence the computed bounds F-statistics and hence the cointegration results, as well as the long-run and short-run results. From the structural break tests, similar conclusion as the standard unit root results follows. The I(1) series are stationary for all variables. As evident from Table 3, the structural breaks in the level series are noted in 2003, 2006 and 2000 for $L_y$, $L_k$ and $LRM$, respectively. In the case of the first difference series, breaks in the series are noted in the years 2003, 2006 and 2000 for $\Delta L_y$, $\Delta L_k$ and $\Delta LRM$, respectively. We factor this information when computing the bounds F-statistics by setting the break period to one in the cumulative dummy ($Dum_{s}$) variable, which is created to account for the structural breaks in the series (see equations (6)-(8)).

The bounds F-statistics are reported in Table 4. The results show evidence of long-run cointegration when the per worker output ($L_y$) is set as the dependent variable. In this case, the computed F-statistics of 10,742.8 exceeds the upper critical bound of 7,783 and 7,607 for the sample size of 35 and 30, respectively, at the 1% level of significance. Notably, when the $L_k$ and $LRM$ are set as dependent variable, separately, the respective computed F-statistics are below the lower critical bounds duly confirming a single cointegrating vector.

After confirming the existence of a long-run relationship between $L_y$, $L_k$, and $LRM$, the diagnostic tests were examined from the ARDL lag estimates. These include: Lagrange multiplier test of residual serial correlation ($\chi^2_{ma}$); Ramsey’s RESET test using the square of the fitted values for correct functional form ($\chi^2_{f}$); normality test based on a test of skewness and kurtosis of residuals ($\chi^2_{n}$); and heteroscedasticity test based on the regression of squared residuals on squared fitted values ($\chi^2_{hs}$). The results are reported in Table 5. In what follows, we find that the diagnostic test rejects the null hypothesis of the presence of serial correlation ($\chi^2_{ma} = 0.0103$), functional form biasness ($\chi^2_{f} = 1.0988$) and heteroscedasticity ($\chi^2_{hs} = 3.8260$) at least at 5% level of significance. However, the test did not reject normality biasness ($\chi^2_{n} = 10.1794$). The presence of non-normality can be attributed to the presence of outliers over the sample period, which results from non-recurring, exogenous shocks (oil price shocks, financial crisis, among other structural shocks) and not the normal evolution of the economic data. This however, can be improved by using pulse dummy variables to capture one-off abnormal observations.

Nevertheless, investigating the CUSUM and CUSUM of squares (CUSUMQ) shows that the parameters of the model are relatively stable over time (Figures 1a and 1b).
Table 5

<table>
<thead>
<tr>
<th>Test Types</th>
<th>LM Version</th>
<th>p-value</th>
<th>F Version</th>
<th>p-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\chi^2_{sc}$</td>
<td>$\chi^2(1) = 0.0103^a$</td>
<td>0.919</td>
<td>F(1, 23) = 0.0076</td>
<td>0.931</td>
</tr>
<tr>
<td>$\chi^2_{ff}$</td>
<td>$\chi^2(1) = 1.0988^a$</td>
<td>0.295</td>
<td>F(1, 23) = 0.8452</td>
<td>0.367</td>
</tr>
<tr>
<td>$\chi^2_{n}$</td>
<td>$\chi^2(2) = 10.1794^a$</td>
<td>0.006</td>
<td>Not applicable</td>
<td></td>
</tr>
<tr>
<td>$\chi^2_{hc}$</td>
<td>$\chi^2(1) = 3.8260^b$</td>
<td>0.050</td>
<td>F(1, 29) = 4.0831</td>
<td>0.053</td>
</tr>
</tbody>
</table>

Notes: A and B indicate rejection of null hypothesis of the presence of the respective Test Types at 1 % and 5 % level of significance, respectively. Source: The authors’ calculation using Microfit 4.1

Short-run and Long-run results

In the long-run (Table 6, Panel a), we note that capital productivity is a dominant driver of growth. The long-run capital share, which is statistically significant at the 1 % level, is 0.50 ($Lk = 0.4999$). As noted, elasticity coefficient of the capital per worker is relatively larger than the stylized value of one-third (0.33) (Rao, 2007; Ertur & Koch, 2007). However, the estimated capital share seems reasonable as over time we expect for some country to realize a higher capital share (Breuss, 2010; Guerriero, 2012). Nevertheless, some reasons for a country to exhibit high capital income shares are when: (a) the capital and labour inputs tend to grow at relatively similar rates; (b) an economy has a large number of self-employed persons who earn income from both capital and their own labour (Gollin, 2002) thus making it difficult to obtain meaningful measures of income shares; and (c) the quality of the data and the sample size which also makes it difficult to compute capital stock (Bosworth & Collins, 2008) that can ideally exhibit decreasing returns to scale and thus conform to a desirable steady-state convergence. Moreover, in the presence of low self-employment rate, the estimated capital share can still be high due to large share of black market (underground) activities in the economy. We therefore concur to all of these reasons in case of this study. The coefficient of remittances (% GDP) ($LRM = 0.0412$) is statistically significant at the 5 % level and contributes about 0.04 % to the long-run output per worker. There is also a marginal positive contribution from the cumulative structural changes that is captured by $Dum_t$ variable. In other words, over the sample period, the structural changes (noted in the years 2000, 2003, 2006, 2007, and 2008) contributed to short-run growth by a factor of 0.11 ($Dum_t = 0.1077$).

The short-run results (Table 6, Panel b) show that output per worker is positively influenced by the lag-one period of the output per worker ($\Delta Ly_t = 0.3608$), that is, previous period short-run growth contributes about 0.36 % to current period output per worker. Moreover, capital per worker, which is positive and statistically significant at the 1 % level, contributes about 1.4% ($\Delta Lk_t = 1.3691$) to output per worker. We also note a marginal positive contribution of remittances in the short-run. In what follows, remittances ($\Delta LRM_t = 0.0194$), which are significant at the 10 % level, contribute about 0.02 % in the short-run. Similar to the positive effects from the long-run results, the structural changes ($Dum_t = 0.0507$) over the sample period contribute to the long-run output by a factor 0.05. The error-correction term ($ECT_{t-1} = -0.4707$), which measures the speed at which prior deviations (errors) from equilibrium are corrected (in this case, about 47 %), has the correct (negative) sign and is significant at the 1 % level duly indicating a relatively speedy convergence to the long-run equilibrium.

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7 Notably, up to 1990 Lithuania was a socialist planning economy where almost all the companies were state-owned. For these years it is not relevant to think in terms of capital income because the wages were arbitrarily administrated.

8 Notably, inter alia, the following mix bag of exogenous shocks is noted here: the 2004 EU membership; in 2000, the dot.com crisis; in 2007, the Schengen agreement membership guaranteeing free labor mobility; 2007–2008, the US financial crisis, which resulted in an economic downturn in 2009.
The Toda-Yamamoto approach to Granger non-causality test

Next, we examine the Granger non-causality test proposed by Toda & Yamamoto (1995) (henceforth T-Y approach). The T-Y approach is suitable when the series are either integrated of different orders, not cointegrated, or both. In these cases, the ECM (error-correction method) cannot be applied for Granger causality tests and the standard (pair-wise) Granger causality test may not give robust results. Hence, the T-Y approach provides a method to test for the presence of non-causality, irrespectively of whether the variables are I(0), I(1) or I(2), not cointegrated or cointegrated of an arbitrary order. In order to carry out the Granger non-causality test, we present the model in the following VAR system:

\[
L_y = \alpha_0 + \sum_{i=1}^{k} \alpha_i L_{y,-i} + \sum_{j=1}^{\text{max}} \eta_j L_{k,-j} + \sum_{j=1}^{\text{max}} \phi_j L_{LM,-j} + \lambda_y
\]

\[
L_k = \beta_0 + \sum_{i=1}^{k} \beta_i L_{k,-i} + \sum_{j=1}^{\text{max}} \theta_j L_{y,-j} + \sum_{j=1}^{\text{max}} \theta_j L_{LM,-j} + \lambda_k
\]

\[
L_{LM} = \gamma_0 + \sum_{j=1}^{\text{max}} \phi_j L_{LM,-j} + \sum_{j=1}^{\text{max}} \phi_j L_{LM,-j} + \sum_{j=1}^{\text{max}} \mu_j L_{k,-j} + \lambda_{LM}
\]

The null hypothesis of non-causality is rejected when the p-values fall within the conventional 1-10% level of significance. Hence, in (9), Granger causality from \(L_k\) to \(L_y\), and \(L_{LM}\) to \(L_y\) implies \(\eta_i \neq 0 \forall i\) and \(\phi_j \neq 0 \forall j\), respectively. Similarly, in (10), \(L_y\), and \(L_{LM}\), granger causes \(L_k\) if \(\beta_i \neq 0 \forall i\), and \(\theta_i \neq 0 \forall i\), respectively; from (11) \(L_{LM}\), \(L_k\), granger causes \(L_{LM}\), if \(\phi_j \neq 0 \forall j\) and \(\mu_i \neq 0 \forall i\), respectively. From the unit root results, the maximum order of integration is 1 (m = 1), and the optimal lag length chosen in the ARDL estimates using the Akaike information (AI) and Schwarz Bayesian (SB) Criteria (Table 6: panel c) is 2 (p = 2). Hence the appropriate lags to carry out the Granger non-causality test (p+m = 3) are 3. Importantly, in conducting the causality tests, it is important to examine the inverse roots of the AR (autoregressive) characteristic polynomial. In order to obtain a robust causality result (based on chi-square and p-values), the inverse roots should lie within the positive and the negative unity. However, where the inverse roots lie outside the unit boundaries, this can be corrected by including appropriate lags, trend and/or the structural break dummy variable as instruments (exogenous variable)
The Granger-causality result (Table 7) shows bidirectional causation between output per worker (Ly) and capital per worker (Lk), which are statistically significant at 1% level. In other words, Ly → Lk (χ² = 53,2577) and Ly ← Lk (χ² = 12,8272) (Ly ←→ Lk) duly indicate the mutually reinforcing effect of the capital and output. Moreover, unidirectional causation is noted from remittances to output per worker (LRM → Ly, χ² = 10,8364) at the 1% level. Finally, the causation from the combined effects (which we define as the causation due conjoint interaction) shows that the capital per worker and the remittances conjointly cause the output per worker (Lk × LRM → Ly, χ² = 15,2472) at the 5% level of significance; and the output per worker and the remittances conjointly causes the capital per worker Ly × LRM → Lk, χ² = 58,58914) at the 1% level of significance.

Conclusion

In this paper, we set out to explore the much controversial topic of whether remittances and hence labor emigration influence economic growth in Lithuania. We used the augmented Solow framework and the ARDL bounds procedure to examine short-run and long run effects, and further extended the study to examine the causality nexus. The results clearly show that the remittances have a dynamic short-run and a long-run momentous effect on the output per worker. Furthermore, the causality results show a mutually reinforcing effect between output per worker and capital per worker, and unidirectional causation from remittances to output per worker. Subsequently, our results support the remittances-led-growth hypothesis (RLG) for Lithuania. Our results in some respects coincide with Elsner (2010) who shows that the workers who stayed in Lithuania have gained from emigration; and with those of Kasnauskiene & Stumbryte (2012) who find that the average contribution of remittances on the economic growth of Lithuania is about 0,0348 %. Moreover, it can be concluded that despite the negative impacts of labor emigration, Lithuanian economy may strongly benefit from labor emigration. Noting these outcomes, a cost effective and efficient remittance inflows to Lithuania need to be encouraged and channeled to ongoing and new productive activities. In the light of an increasing share of emigrating families and Lithuania’s demographic developments, it may be appropriate to consider introducing incentives schemes to make re-migration more attractive.

Some caveats to the results are in order. It is important to point out that our results are not unambiguous due to data constraints and the approximations of the data points to fill the missing data. To address this issue within the limits, we assumed a priori cointegration and examined short-run and long-run coefficients of capital stock per worker and remittances and the causality effects with the sample size of 19 data points (1993–2011) from the World Bank (2013) database with lag-length of 2 and 1 to estimate the ARDL results’ causality, respectively. Due to space limitation and the limitations of small sample size, we do not report the full results of here. However, we briefly discuss them at this point to trigger further discussion. Cutting the sample size to the period of 1993–2011, we find that the long-run coefficient of capital per worker is 0,68, which is relatively high. Although the relatively capital share is plausible, the complexity lies in re-estimating the initial capital stock and selecting the ‘appropriate’ depreciation rate which may give varied results. For instance, instead of cutting the data from the original sample size of the period of 1980–2012, we can re-estimate the capital stock with the sample size of the period of 1993–2011 by computing initial capital and assuming appropriate depreciation rate. With a small sample size, computing the capital stock also becomes challenging and sometimes may require the use of high depreciation rate, which may be difficult to justify. In terms of the coefficient of remittances, it is found that remittances, at best, are still positive (both in the short-run (0,013 %) and the long-run (0,018 %), however, are not statistically significant within the 1–10 percent conventional level of statistical significance. In regards to the causality procedure, we find an indirect causation from capital stock to remittances (at 10 % level of statistical significance) duly indicating that remittances Granger cause capital investment, and from output to capital stock (at 1 percent level of statistical significance). However, another set of caveat lies in case of the latter estimation. Notably, using of the actual data with the bounds procedure is not appropriate since the procedure is developed for the sample size of at least 80, and for sample size between 30 and 80, one has to either re-compute the bounds or use the Narayan (2005) critical bounds to examine the cointegration. Moreover, for a relatively small sample size of say 19, as in this case, there will be required specific computation of the appropriate bounds to examine the cointegration and the short-run and the long-run results. Therefore, further research with appropriate method of analysis may support (or otherwise) the results presented here, and subsequently trigger more discussion on the approaches used and the broad-based role of remittances in the Lithuanian economy.

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Appendix

### Table A

#### Summary of data compilation method

<table>
<thead>
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<th>Variable</th>
<th>Years available in the WDI</th>
<th>Method/function used for approximation</th>
<th>Period for which data is approximated</th>
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<tr>
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<td>1990–2012</td>
<td>Exponential: $y = 5E+09e^{0.1054x}$</td>
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<td>Gross fixed capital formation (Constant 2005 USD)</td>
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<td>Exponential: $y = 3E+09e^{0.0602x}$</td>
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<td>Personal Remittances received (Current Prices)</td>
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### Table B

#### Sample used for Lithuania (1980-2012)

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<th>GDP (current USD)</th>
<th>GDP (constant 2005 USD)</th>
<th>Investment (constant 2005 USD)</th>
<th>Remittances (Current USD)</th>
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NB: Data on employment is computed from the average employment rate. Missing data was approximated using the method in Table A. Source: World Bank (2013)
Reference


