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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).

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

[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 .

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, Krupickaite & Poviliunas, 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 compliment 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.

$y_t = A_t k_t^{\alpha}$, α > 0 (1)

where A = stock of technology and k = capital per worker, and α is the capital share. The Solow model assumes that the evolution of technology is given by:

$\Phi_t = A_o e^{gt}$ (2)

where A_0 is the initial stock of knowledge and t is time.

Next, we introduce remittances, RM_t, as shift variable:

$\Psi_t = f(RM_t)$ (3)

The effect of RM_t on total factor of 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^{β} (where $\beta \in [0, 1]$ represents elasticity of remittances) as part of the stock of technology and redefine A_t as follows:

$A_t = \Phi_t \Psi_t = A_o e^{gt} RM_t^{\beta}$ (4)

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

Hence,

$y_t = (A_o e^{gt} RM_t^{\beta}) k_t^{\alpha}$ (5)

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 (δ) 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 the average rate of employment (which was 52,3 % over the period 1991–2011). Remittances data are expressed as percent of current GDP. A total of 33 years of annual data over the period of 1980–2012 are used in the analysis. A summary of the data compilation method is provided in Table A and the sample data in Table B (Appendix). All data on the key variables are sourced from *World Development Indicators and Global Development Finance* database (World Bank, 2013). All data are duly transformed into natural log form for analysis. Hence, *Ly* = natural log of output per worker (*y*); *Lk* = natural log of capital per worker (*k*), and *LRM* = natural log of remittances as a percent of GDP. A descriptive statistics and correlation matrix of log of output and capital per worker, and remittances (as percent of GDP) are provided in Table 1.

Table 1

Descriptive statistics and correlation matrix

	Ly	*Lk*	*LRM*
Mean	9,0135	9,3587	-4,3197
Median	8,9887	9,4392	-4,0715
Maximum	9,8649	10,649	1,9904
Minimum	8,2906	7,5975	-12,1956
Std. Dev.	0,5312	0,8214	4,6779
Skewness	0,1555	-0,3425	-0,0681
Kurtosis	1,6052	2,2007	1,6382
Jarque-Bera	2,8079	1,5237	2,5756
Probability	0,2456	0,4668	0,2759
Ly	1,0000	-	-
Lk	0,9688	1,0000	-
LRM	0,9836	0,9799	1.0000

Source: the authors' calculations using Eviews 8

ARDL bounds procedure

Next, we specify the ARDL specifications as (6)-(8) below. Note that each equation has a dummy variable (Dum_s) 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_s therefore provides a relatively more robust computation of bound statistics and also short-run and long-run results.

$$\Delta Ly_t = \beta_{10} + \beta_{11} Ly_{t-1} + \beta_{12} Lk_{t-1} + \beta_{13} LRM_{t-1} + \alpha_{10} Dum_s + \sum_{i=1}^{p} \alpha_{11i} \Delta Ly_{t-i} + \sum_{i=0}^{p} \alpha_{12i} \Delta Lk_{t-i} + \sum_{i=0}^{p} \alpha_{13i} \Delta LRM_{t-i} + \varepsilon_{1t} \quad (6)$$

$$\Delta Lk_t = \beta_{20} + \beta_{21} Ly_{t-1} + \beta_{22} Lk_{t-1} + \beta_{23} LRM_{t-1} + \alpha_{20} Dum_s + \sum_{i=1}^{p} \alpha_{21i} \Delta Ly_{t-i} + \sum_{i=0}^{p} \alpha_{22i} \Delta Lk_{t-i} + \sum_{i=0}^{p} \alpha_{23i} \Delta LRM_{t-i} + \varepsilon_{2t} \quad (7)$$

$$\Delta LRM_t = \beta_{30} + \beta_{31} Ly_{t-1} + \beta_{32} Lk_{t-1} + \beta_{33} LRM_{t-1} + \alpha_{30} Dum_s + \sum_{i=1}^{p} \alpha_{31i} \Delta Ly_{t-i} + \sum_{i=0}^{p} \alpha_{32i} \Delta Lk_{t-i} + \sum_{i=0}^{p} \alpha_{33i} \Delta LRM_{t-i} + \varepsilon_{3t} \quad (8)$$

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 *el 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 emphasize the 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; and 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.

Table 2

Unit root tests results

Variables in log form	ADF		Phillips and Perron		KPSS	
	Level	1st Diff.	Level	1st Diff.	Level	1st Diff.
Ly_t	-2,6183	-4,9003^A	-2,6026	-4,4224^A	0,1239^C	0,1740^A
Lk_t	-14,3885^A	-5,7153^A	-14,3885^A	-9,2908^A	0,1781^C	0,4665^B
LRM	-3,4697^B	-8,0253^A	-3,4406^B	-12,8731^A	0,0949^A	0,4285^B

NB: The ADF critical values are based on MacKinnon (1996). The optimal lag is chosen on the basis of the Akaike Information Criterion (AIC). The null hypotheses for ADF and Phillips-Perron (KPSS) tests are that a series has a unit root or is non-stationary (stationary), respectively. A, B and C denote 1 %, 5 % and 10% level of significance at which the null hypothesis is rejected (and accepted) in case of the ADF and Phillips-Perron (KPSS) tests. Source: The authors' calculation using Eviews 8

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, 2007 and 2000 for *Ly*, *Lk* 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 ΔLy, ΔLk and ΔLRM, 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)).[3]

Table 3

Unit root tests with single structural break

Variables	Level		1st Diff.	
	T-stat	Break	T-stat	Break
Ly	-3,437990[1]	2003	-6,105669[1]^A	2003
Lk	-5,518199[1]^B	2007	-6,884112[1]^A	2006
LRM	-5,885048 [0]^A	2000	-8,942652[0]^A	2000

NB: Critical values are obtained from Zivot & Andrews (2002). The null hypothesis is that a series has a unit root with a structural break in both the intercept and the trend; A denotes rejection of the null hypothesis at the 1 % level of significance. Source: The authors' calculation using Eviews 8

Table 4

Results of Bound Tests based on 1 % critical bounds

Dependent Variable		Computed F-statistic
Ly		10,7428^A
Lk		2,5870
LRM		2,8600
Sample size	Lower bound value	Upper bound value
35	6,183	7,873
30	6,140	7,607

NB: Critical values are from Narayan (2005) - Critical values for the bounds test: case V: unrestricted intercept and no trend, p. 1988; Dum_s = 1 for 2000, 2003, 2006, 2007, and 2008. A - indicates significance at 1% level. Source: The authors' calculation using Microfit 4.1

The bounds F-statistics are reported in Table 4. The results show evidence of long-run cointegration when the per worker output (*Ly*) is set as the dependent variable. In this case, the computed F-statistics of 10,7428 exceeds the upper critical bound of 7,873 and 7,607 for the sample size of 35 and 30, respectively, at the 1 % level of significance. Notably, when the *Lk* 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.[4]

After confirming the existence of a long-run relationship between Ly_t, Lk_t, and LRM_t, the diagnostic tests were examined from the ARDL lag estimates.[5] These include: Lagrange multiplier test of residual serial correlation (χ^2_{sc}); Ramsey's RESET test using the square of the fitted values for correct functional form (χ^2_{ff}); normality test based on a test of skewness and kurtosis of residuals (χ^2_n); and heteroscedacity test based on the regression of squared residuals on squared fitted values (χ^2_{hc}). 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 (χ^2_{sc} = 0,0103), functional form biasness (χ^2_{ff} = 1.0988) and heterocedacity (χ^2_{hc} = 3,8260) at least at 5 % level of significance. However, the test did not reject normality biasness (χ^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.[6] 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).

[3] Moreover, including 2008 in the structural break to account for the financial crisis as further improved the results all across.

[4] Note that the critical bounds are presented for the samples of 35 and 30 since the Narayan (2005) bounds are for the samples from 30 to 80 with 5-year intervals. The sample size in this study is 33 and the computed F-statistics of *Ly* as dependent variable clearly exceeds both samples of 30 and 35.

[5] The ARDL lag estimation results are not included here to conserve space. We only provide the diagnostic tests in order to ascertain the robustness of the long-run and short-run results.

[6] Although we understand that including the pulse dummy variables will improve the diagnostic test results, we leave this for further research.

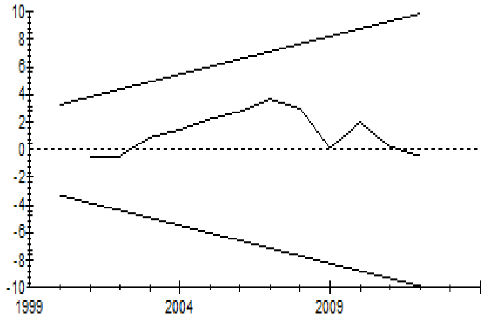

Figure 1a. Cumulative Sum of Recursive Residuals
Note: The straight lines represent critical bounds at 5 % significance level

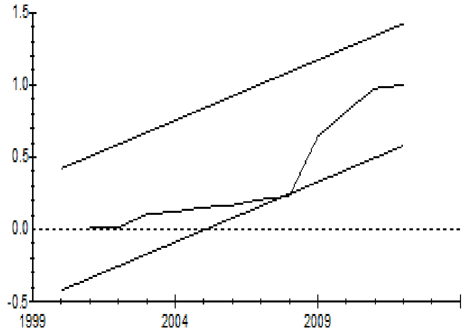

Figure 1b. Cumulative Sum of Squares of Recursive Residuals
Note: The straight lines represent critical bounds at 5 % significance level

Table 5

Diagnostic tests from the ARDL lag estimates

Test Types	LM Version	p-value	F Version	p-value
χ^2_{sc}	$\chi^2(1) = 0{,}0103^A$	0,919	$F(1, 23) = 0{,}0076^A$	0,931
χ^2_{ff}	$\chi^2(1) = 1{,}0988^A$	0,295	$F(1, 23) = 0{,}8452^A$	0,367
χ^2_{n}	$\chi^2(2) = 10{,}1794$	0,006	Not applicable	
χ^2_{hc}	$\chi^2(1) = 3{,}8260^B$	0,050	$F(1, 29) = 4{,}0831^B$	0,053

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[7]; and (c) the quality of the data and the sample sizem 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_s variable.[8] 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_s = 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 (ΔLy_{t-1} = 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% (ΔLk_t = 1,3691) to output per worker. We also note a marginal positive contribution of remittances in the short-run. In what follows, remittances (ΔLRM_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_s = 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.

[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.

Table 6

Estimated long run coefficients and error correction representation

Panel a: Long-run (dependent variable Ly_t)				*Panel b: Short-run (dependent variable ΔLy_t)*			
Regressor	Coefficient	Standard Error	t-ratio	Regressor	Coefficient	Standard Error	t-ratio
Lk	0,4999	0,1168	4,2790^A	ΔLy_{t-1}	0,3608	0,1469	2,4559^B
LRM	0,0412	0,0194	2,1186^B	ΔLk_t	1,3691	0,3743	3,6581^A
Dum_s	0,1077	0,0473	2,2768^B	ΔLRM_t	0,0194	0,0101	1,9140^C
Constant	4,2708	1,1931	3,5797^A	Dum_s	0,0507	0,0225	2,2526^B
				Constant	2,0104	0,7119	2,8239^A
				ECT_{t-1}	-0,4707	0,0924	-5,0951^A

Panel c: Short-run dynamics test statistics			
R-Squared	0,5559	R-Bar-Squared	0,4449
S.E. of Regression	0,0368	F-stat. F(5, 25)	6,0080
Mean of Dependent Variable	0,0503	S.D. of Dependent Variable	0,0494
Residual Sum of Squares	0,0325	Equation Log-likelihood	62,3647
Akaike Info. Criterion	55,3647	Schwarz Bayesian Criterion	50,3457
DW-statistic	1,9624	ARDL(2, 1, 0)	N = 33

Notes: A, B, and C refer to the 1%, 5%; and 10% level of significance, respectively; Source: The authors' calculation using Microfit 4.1.

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:

$$Ly_t = \alpha_0 + \sum_{i=1}^{k}\alpha_{1i}ly_{t-i} + \sum_{j=k+1}^{d\max}\alpha_{2j}Ly_{t-j} + \sum_{i=1}^{k}\eta_{1i}Lk_{t-i} + \sum_{j=k+1}^{d\max}\eta_{2j}Lk_{t-j} + \sum_{i=1}^{k}\phi_{1i}LRM_{t-i} + \sum_{j=k+1}^{d\max}\phi_{2j}LRM_{t-j} + \lambda_{1t} \tag{9}$$

$$Lk_t = \beta_0 + \sum_{i=1}^{k}\beta_{1i}Lk_{t-i} + \sum_{j=k+1}^{d\max}\beta_{2j}Lk_{t-j} + \sum_{i=1}^{k}\theta_{1i}ly_{t-i} + \sum_{j=k+1}^{d\max}\theta_{2j}Ly_{t-j} + \sum_{i=1}^{k}\vartheta_{1i}LRM_{t-i} + \sum_{j=k+1}^{d\max}\vartheta_{2j}LRM_{t-j} + \lambda_{2t} \tag{10}$$

$$LRM_t = \gamma_0 + \sum_{i=1}^{k}\gamma_{1i}LRM_{t-i} + \sum_{j=k+1}^{d\max}\gamma_{2j}LRM_{t-j} + \sum_{i=1}^{k}\varphi_{1i}Ly_{t-i} + \sum_{j=k+1}^{d\max}\varphi_{2j}Ly_{t-j} + \sum_{i=1}^{k}\mu_{1i}Lk_{t-i} + \sum_{j=k+1}^{d\max}\mu_{2j}Lk_{t-j} + \lambda_{3t} \tag{11}$$

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 Lk_t to Ly_t, and LRM_t to Ly_t, implies $\eta_{1i} \neq 0 \forall i$ and $\phi_{1i} \neq 0 \forall i$, respectively. Similarly, in (10), Ly_t, and LRM_t granger causes Lk_t if $\theta_{1i} \neq 0 \forall i$, and $\vartheta_{1i} \neq 0 \forall i$, respectively; from (11) Ly_t, Lk_t granger causes LRM_t if $\varphi_{1i} \neq 0 \forall i$ and $\mu_{1i} \neq 0 \forall i$, 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 (auto-regressive) characteristics 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) in the VAR equation. We ensured the AR inverse roots are within the positive/negative unit boundary by including the cumulating structural breaks (Dum_s) as exogenous variable before proceeding to the causality assessment. The results of the causality tests are reported in Table 7.

Table 7

Granger non-causality test based on χ^2 (p-value)

		Dependent Variable (**Y**)		
		Ly	*Lk*	*LRM*
Excluded variables (**X**) →	*Ly*	-	53,2577^A (0,0000)	1,5302 (0,6753)
	Lk	12,8272^A (0,0050)	-	1,6464 (0,6489)
	LRM	10,8364^A (0,0126)	2.8137 (0.4213)	-
	Combined	15,2472^B (0,0184)	58,58914^A (0,0000)	3,5702 (0,7346)

Notes: Based on the maximum order of integration (p=1) and maximum lag used in the ARDL estimation (m=2), the maximum lag in the Toda-Yamamoto non-causality test is (p+m) 3. A and B indicate causality at the 1 % and 5 % levels of significance, respectively. Granger-causality is denoted as X causing Y, that is: X→Y. *Source*: The authors' calculation using Eviews 8.

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* (χ^2 = 53,2577) and *Ly* ← *Lk* (χ^2 = 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*, χ^2 = 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*, χ^2 = 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*, χ^2 = 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 depreciate 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

Variable	Years available in the WDI	Method/function used for approximation	Period for which data is approximated
GDP (Current Prices)	1990–2012	Exponential: $y = 5E{+}09e^{0.1054x}$ $R^2 = 0{,}9346$	1980–1989
Gross fixed capital formation (Constant 2005 USD)	1995–2012	Exponential: $y = 3E{+}09e^{0.0602x}$ $R^2 = 0{,}6877$	1980–1994
Personal Remittances received (Current Prices)	1993–2011	Exponential: $y = 177202e^{05636x}$ $R^2 = 0{,}9238$	1980–1992 and 2012

Table B

Sample used for Lithuania (1980-2012)

Year	GDP (current USD)	GDP (constant 2005 USD)	Investment (constant 2005 USD)	Remittances (Current USD)	Population	Employment
1980	1936414200	7120000000	1291510274	360	3413202	1786080
1981	2151656289	7280000000	1371647111	632	3432947	1796412
1982	2390823608	7480000000	1456756354	1111	3457179	1809092
1983	2656575567	7720000000	1547146534	1951	3485192	1823751
1984	2951867180	8000000000	1643145329	3428	3514205	1838933
1985	3279981927	8320000000	1745100747	6024	3544543	1854809
1986	3644568264	8680000000	1853382390	10583	3578914	1872795
1987	4049680189	9080000000	1968382794	18595	3616367	1892393
1988	4499822324	9520000000	2090518850	32671	3655049	1912635
1989	5000000000	10000000000	2220233318	57403	3684255	1927918
1990	5555774917	10520000000	2357996431	100856	3697838	1935026
1991	6173326986	11080000000	2504307599	177202	3704134	1938320
1992	6859523046	11680000000	2659697218	311341	3700114	1936217
1993	7621993216	12320000000	2824728599	34530	3682613	1927059
1994	6958636950	13898969284	3000000000	743072	3657144	1913731
1995	7904895791	14356256518	3186146805	1095000	3629102	1899057
1996	8426600000	15100293918	3383843822	2575000	3601613	1884673
1997	10128700000	16228155777	3593807727	2927500	3575137	1870818
1998	11254050000	17466172080	3816799670	3425000	3549331	1857314
1999	10971375000	17278719288	4053628025	2995000	3524238	1844183
2000	11434200000	17840402458	4305151327	49942501	3499536	1831257
2001	12159225000	19042080928	4572281382	79150002	3481292	1821710
2002	14163949142	20349038400	4855986573	109248047	3469070	1815315
2003	18608709857	22434139923	5157295368	114834000	3454205	1807536
2004	22551543054	24083224893	5477300054	324496735	3435591	1797796
2005	25962254181	25962254181	5817160690	534305077	3414304	1786657
2006	30088510798	27998948427	6178109316	994085129	3394082	1776075
2007	39103973051	30753989500	6561454419	1432637624	3375618	1766413
2008	47252926429	31654296590	6968585679	1565807821	3358115	1757254
2009	36846183172	26987911154	7400978999	1239407206	3339456	1747490
2010	36644563174	27390349481	7860201865	1673614993	3286820	1719946
2011	42890914111	29022647091	8347919020	1956390868	3030173	1585646
2012	42086555101	30062313805	8865898506	2200000000	2985509	1562274

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)

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Perlaidų įtakos Lietuvos ekonominiam augimui tyrimas

Santrauka

Perlaidos gali daryti teigiamą, arba neigiamą įtaką perlaidas gaunančios šalies augimui, priklausomai nuo to ar perlaidos panaudojamos vartojimo, ar investavimo tikslams. Perlaidų srautas į gaunančias šalis labiausiai priklauso nuo darbuotojų, kurie emigravo, bet vis dar turi glaudų ryšį su namais. Darbuotojų emigracija neabejotinai mažina darbo jėgą ir žmogiškąjį kapitalą gimtojoje šalyje, apie kurią dažnai kalbama kaip apie pradedančią ilgą, mažesnio BVP augimo tempo laikotarpį. Tačiau, jei emigravę darbuotojai persiunčia žymią užsienyje uždirbtų pajamų dalį ir jei tas perlaidas gaunantys giminės jas panaudoja vartojimo tikslams, tada, geriausiu atveju, perlaidos nedaro arba daro nežymią įtaką ekonominiam augimui per vartojimo kanalą. Blogiausiu atveju, perlaidos sukelia ryškų vartojimą (Rempel ir Loddell, 1978), kuris gali erzinti ne emigrantus. Perlaidos gali varžyti perlaidas gaunančių migrantų, giminių darbo jėgos pasiūlą gimtojoje šalyje. Taigi, abiejų atvejų įtaka (migrantai, siunčiantys perlaidas ir vietinės darbo jėgos pasiūlos mažėjimas), ilgainiui gali sumažinti augimo tempą. Perlaidų egzistavimas taip pat gali sukurti moralinį pavojų perlaidas gaunantiems giminaičiams gimtojoje šalyje (Chami *ir kt*., 2005) ir/arba gimtosios šalies valdžiai (Grabel, 2009; Singer, 2012). Abiem atvejais, perlaidos gali sukelti gimtosios šalies veiksnių ekonominių pastangų mažėjimą, kurio pasekmės būtų tai, kad augimo tempas taps mažesnis nei įmanoma. Perlaidas gaunančių giminių atžvilgiu, migrantas yra tarsi draudimas nuo ekonominės rizikos gimtojoje šalyje. Valdžios atžvilgiu, perlaidų egzistavimas gali paskatinti didelį skolos ir BVP santykį. Abiejų rūšių moralinis pavojus sumažina augimo tempą dėl mažesnių vietinių santaupų ir didesnių palūkanų išmokų.

Iš kitos pusės, kai savo gimtojoje šalyje darbuotojai yra bedarbiai, tokių darbuotojų emigracija sumažina valdžios vartojimo išlaidas, tokiu būdu sumažina ir likusių piliečių mokesčių naštą, vėliau sustiprina ekonomikos augimo tikimybę. Be to, jei pervesti uždarbiai yra siunčiami namo ir panaudojami investavimui gimtojoje šalyje, mes galime tikėtis teigiamos augimo įtakos (Rapoport, 2002; Rapoport ir Doquier, 2006; Stark, 1991). Taip pat, galimybė gauti pajamas užsienyje, gali sukelti didesnį investavimą į žmogiškąjį kapitalą, kas toliau didins migrantų darbuotojų galimybes įsidarbinti užsienyje. Abiejų rūšių investicijos tol skatina pajamų augimo tempą gimtojoje šalyje, kol darbuotojas perveda dalį savo pajamų namo. Manant, kad dauguma migrantų yra kilę iš mažas arba vidutines pajamas gaunančios klasės, perlaidos taip pat gali padėti įveikti likvidumo suvaržymus, kurie, kitu atveju, dominuotų ir tuo pat metu mažintų turtinę nelygybę gimtojoje šalyje (Stark *ir kt*., 1986, Taylor ir Wyatt, 1996). Tačiau, jei tokios prielaidos nėra ir, kraštutiniu atveju, jei migracijos kaštai yra palyginti aukšti, gali būti išprovokuota priešinga įtaka turto paskirstymui (Rapoport, 2002). Todėl nestebina, kad politikai, žiniasklaida, tyrėjai ir visuomenė, yra šiek tiek neapsisprendę dėl tarptautinės darbo jėgos mobilumo teigiamos įtakos, o ypač emigracijos atveju. Lietuva buvo viena iš greičiausiai augančių ekonomikų Europos Sąjungoje nuo tada, kai tapo Europos Sąjungos nare, iki 2008 metų finansinės krizės. Dabar Lietuva laikoma kaip didelių pajamų šalis ir todėl daugiau nebėra įtraukta į besivystančių, perlaidas gaunančių šalių grupę. Apibūdinama kaip grynos emigracijos šalis, ji susiduria su įvairiais iššūkiais ir nauda. Teigiami dalykai tie, kad kiekvienais metais šalis užregistruoja didelius kiekius perlaidų. Tačiau, kai kurie tyrimai pateikė skirtingus rezultatus apie perlaidas į Lietuvą (Kasnauskienė ir Buzytė, 2011; Damulienė, 2013; Kasnauskienė ir Stumbrytė, 2012). Galiausiai, bendra perlaidų įtaka augimui nėra išskirtinė. Ji priklauso nuo siunčiančios šalies ekonominės situacijos, kultūros ir institucijų. Šį tyrimą pirmiausiai paskatino faktas, kad ekonometrinis perlaidų įtakos vieno darbuotojo išdirbiui tyrimas, davė skirtingus rezultatus. Tai nestebina, nes dauguma tyrimų sutelkia dėmesį į šalis su didesniais emigracijos tempais, kurios paprastai būna mažiau išsivysčiusių šalių grupėje. Lietuva yra viena iš Europos Sąjungos šalių-narių, kurios nuo savo nepriklausomybės realizavo pastovią *neigiamos migracijos normą* ir užregistravo vieną iš didžiausių *neigiamos grynosios migracijos normų*. Be to, Lietuvos gyventojų etninė sudėtis yra homogeniškesnė nei Latvijos ir Estijos, nors Lietuvos gyventojų skaičius yra tik tris kartus didesnis už Estijos. Lietuvai emigracija daro neigiamą įtaką. Pirmiausia, Lietuvoje mažas gimstamumas, todėl Lietuva yra viena iš greičiausiai senėjančių visuomenių Europoje, o. antra, dauguma emigrantų yra jauni ir gerai išsilavinę.

Straipsnyje analizuojamas Lietuvoje gaunamų perlaidų vaidmuo atsižvelgiant į tai, kad šia tema yra paskelbta mažai darbų. Perlaidų vaidmuo Lietuvoje nagrinėtas kitokiu aspektu. Šiame darbe tiriama perlaidų įtaką vieno darbuotojo išdirbiui Lietuvoje per pasirinktą 1980–2012 metų laikotarpį. Taikytas papildytas Solow metodas ir ARPV (autoregresinis paskirstytas vėlavimas), kad būtų išnagrinėtas kointegracijos ryšys ir vėliau įvertinta ilgalaikė ir trumpalaikė įtaka bei priežastinis priklausomumas. Prieinamų *Pasaulio Banko* duomenų, siekiant gauti informacijos, reikalingos kintamiesiems nustatyti, yra gan mažai. Šiai kliūčiai pašalinti, mes siūlome apytikslį metodą, kaip surinkti trūkstamus duomenis. Asmeninių perlaidų atveju, mes naudojame eksponentines krypties funkcijas, kaip tinkamiausias, norint priartinti perlaidų duomenis per 1980–1992 ir 2012 metų laikotarpį. Panašus metodas naudojamas norint priartinti BVP duomenis (dabartinėmis kainomis) per 1980–1989 metų laikotarpį, taip pat suformuoti bendrojo pagrindinio kapitalo duomenis apie akcinį kapitalą per 1980–1994 metų laikotarpį. Taip pat naudojama polinominė krypties funkcija, kad būtų priartinti BVP duomenys (konstanta 2005 USD) 1980–1994 metų laikotarpiui. Vėliau, rezultatai rodo, kad vieno darbuotojo išdirbis, kapitalas vienam darbuotojui ir perlaidos yra kointegruotos. Perlaidos prisideda maždaug 0,02 % ir 0,04 % prie vieno darbuotojo išdirbio per ilgą ir trumpą laikotarpį. Aišku, ši ilgo laikotarpio perlaidų dalis sutampa su Kasnauskienės ir Stumbrytės (2012) darbu. Tačiau mes pastebime dvikryptį priežastingumą, taigi ir abipusiai stiprinantį efektą, tarp kapitalo vienam darbuotojui ir vieno darbuotojo išdirbio, ir vienakryptį priežastingumą iš perlaidų į vieno darbuotojo išdirbį. Rezultatai nėra vienareikšmiai dėl duomenų apribojimų ir apytikrio duomenų kiekio, kad užpildytų trūkstamus duomenis. Siekiant išspręsti šį klausimą, ėmėmės *a priori* kointegracijos ir ištyrėme trumpo ir ilgo laikotarpio akcinio kapitalo vienam darbuotojui koeficientus ir perlaidas bei priežastingumo įtaką pavyzdžiu, kurį sudarė 19 duomenų taškų (1993–2011) iš *Pasaulio Banko* (2013) duomenų bazės su vėlavimu 2 ir 1, kad atitinkamai įvertintume ARPV rezultatų priežastingumą. Dėl erdvės apribojimų ir mažo dydžio pavyzdžio apribojimų, nepateikti išsamūs rezultatai. Jie aptarti trumpai, kad būtų sukeltos tolesnes diskusijos. Taigi, sumažinus *pavyzdžio dydį* iki 1993–2011 metų, mes sužinome, kad ilgo laikotarpio kapitalo, vienam darbuotojui, koeficientas yra 0,68, kuris yra pakankamai aukštas. Nors santykinė kapitalo dalis yra patikima, tačiau problema yra pakartotinis pradinio akcinio kapitalo įvertinimas ir „atitinkamo" nuvertėjimo koeficiento pasirinkimas, nes jie gali duoti kintančius rezultatus. Pavyzdžiui, vietoj to, kad sumažintume duomenų kiekį iš originalaus 1980–2012 metų pavyzdžio, mes galime pakartotinai įvertinti akcinį kapitalą, panaudodami 1993–2011 metų pavyzdį, apskaičiuodami pradinį kapitalą ir numanydami atitinkamą nuvertėjimo koeficientą. Esant *mažam pavyzdžiui*, akcinio kapitalo apskaičiavimas taip pat tampa iššūkiu, o kartais gali pareikalauti panaudoti aukštą nuvertėjimo koeficientą, kurį pateisinti būtų sunku. Remiantis perlaidų koeficientu, nustatyta, kad perlaidos, geriausiu atveju, vis dar yra teigiamos (ir per ilgą (0,013 %), ir per trumpą laikotarpį (0,018 %)), tačiau jos nėra statistiškai reikšmingos, esant 1–10 procentų įprastiniam statistinio reikšmingumo lygiui. *Priežastingumo procedūros* atžvilgiu, mes randame netiesioginį priežastingumą iš akcinio kapitalo į perlaidas (esant 10 % statistinio reikšmingumo lygiui), kuris tinkamai parodo, kad perlaidos *Granger priežastingumas* sukelia kapitalo investicijas iš išdirbio į akcinį kapitalą (esant 1 procento statistinio reikšmingumo lygiui). Tačiau, kitas įspėjimų rinkinys slypi paskutinio įvertinimo atveju. Aišku, naudoti tikrus duomenis su ribų procedūra netinka, nes procedūra yra sukurta, kurios dydis yra mažiausiai 80, ir pavyzdžiams, kurių dydis tarp 30 ir 80, reikėtų arba perskaičiuoti ribas, arba naudoti Narayan (2005) kritines ribas, norint išnagrinėti kointegraciją. Taigi *mažam pavyzdžiui*, kurio dydis, tarkime, 19 (kaip anksčiau aptartu atveju), reikės tam tikro atitinkamų ribų apskaičiavimo, norint išnagrinėti kointegraciją bei ilgo ir trumpo laikotarpių rezultatus. Todėl tolesnis tyrimas, su tam tikru analizės metodu, gali patvirtinti (arba ne), čia pateiktus rezultatus, o vėliau paskatinti daugiau diskutuoti apie naudotus metodus ir plačiai pagrįstą perlaidų vaidmenį Lietuvos ekonomikoje. Galiausiai, šiame darbe pateikiama alternatyvi struktūra ir metodas, kaip įvertinti ilgo ir trumpo laikotarpio perlaidų įtaką ir priežastingumą, kadangi paskutiniuoju atveju, teigiamos koreliacijos nebūtinai pakanka priežastingumui. Tačiau, empiriniai rezultatai gali būti prieštaringi ir todėl paneigiami esant įspėjimams, kurių ribose atliekama analizė. Šia pastaba mes skatiname tolesnius debatus, svarstymus ir diskusijas, o esant patikimesniems bei vienodiems duomenims, sutinkame, kad panaudojus šią struktūrą ir metodą, rezultatus galima pagerinti.

Raktažodžiai: *perlaidos, kapitalas vienam darbuotojui, kointegracija, ARPV metodas, Granger priežastingumo testai, ekonominis augimas, Lietuva.*

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