DOES PUBLIC DEBT AFFECT ECONOMIC GROWTH? PANEL EVIDENCE FROM CENTRAL AND EASTERN EUROPE

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Abstract
The paper employs a form of the panel ARDL-based error correction model (ECM) to explore the short-run and long-run relationship between public debt (and certain macroeconomic variables) and economic growth in Central and Eastern European (CEE) countries. It covers the period 2006Q1−2018Q4. The results indicate both short-term and long-term causality (except for financial development), though the marginal effects are variable-specific. There is a negative effect of public debt, interest rate and exchange rate on growth, whereas the saving rate, trade openness, financial development, fixed capital formation and population growth contribute to economic development. Thus, a responsible and saving-oriented fiscal policy, coupled with higher private sector investment, an export-oriented private sector and undervalued real exchange rate, and population growth would contribute to economic development in CEE countries.

Key words: Public debt, economic growth, CEE countries, ARDL-EC Model
JEL Classification: E62, C33, H63, O52

Introduction

The relationship between public debt and economic activity, coupled with the potential economic and political risks of high public debt, has attracted considerable attention of both policymakers and politicians. Opinions regarding this relationship have varied significantly over time and modern economic theory does not have an unambiguous standpoint about the economic implications of public debt. Some of the most influential schools of economic thought have held diametrically opposite views on public debt (Salsman, 2017). For example, the Keynesian concept, as a reaction to the previous

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viewpoint of a balanced government budget, is based on the belief that high public debt is a government asset rather than a liability. It is also based on the concept that a government running a deficit encourages economic activity in the long run. In contrast, the neoclassical school advocates against public debt. It is argued that high public spending is financed through borrowing, and it harms rather than promotes economic growth since savings are not distributed to productive investments. Furthermore, private investment is crowded out, which in turn undermines capital and wealth accumulation.

Currently, many countries, both developing and developed, are struggling with high public debts, budget deficits and sluggish economic growth. Therefore, the main question researchers and policymakers are faced with is related to the transmission channels and magnitude of the public debt effects on a country's macroeconomic performance. The problem has been fuelling a new academic debate about the interaction between public debt and macroeconomic performance, primarily the debt-growth causality. Some studies show that public debt has a negative impact on economic growth (see, *e.g.*, Čeh-Časni *et al.*, 2014; Bokemeier and Greiner, 2015; Ferreira, 2016), while others reveal a positive impact (see, *e.g.*, Fincke and Greiner, 2015). Finally, some other studies find no evidence that high levels of public debt undermine economic growth (Panizza and Presbitero, 2014). However, most studies show that the relationship between public debt and economic growth is non-linear and concave. Accordingly, low levels of public debt stimulate economic growth, while high levels of public debt, beyond a specific level relative to the GDP, are associated with lower and even negative economic growth rates, in both developed and developing countries (Reinhart and Rogoff, 2010; Minea and Parent, 2012; Mencinger *et al.,* 2014; Alfonso and Alves, 2015; Dincă and Dincă, 2015; Woo and Kumar, 2015). A potential explanation is that a country usually needs to borrow in order to accelerate its economic growth, but after a certain threshold is reached, an additional unit of debt discourages investment, causing the growth rate to decline over time.

Central and Eastern European (CEE) countries have undergone a painful process of both economic and political transition; there are also some adverse effects of the global economic crisis of 2007–2009, with severe spillover effects throughout the region. To alleviate this situation, the governments have implemented numerous fiscal reforms, though public debt remained high. The public debt crisis and the consequent economic slowdown were accelerated by inefficient economic systems and slow transition processes, *i.e.*, low productivity reflecting low investment rates, weak institutions and a difficult business environment (World Bank, 2017). Accordingly, the main goal of this study is to examine the relationship between public debt and economic growth in CEE countries, since this relationship has significant policy implications. More specifically, we formulate three research questions: (1) What is the relationship between public debt and growth?

(2) Is the relationship non-linear or linear? (3) How much do public debt fluctuations determine the fluctuations in GDP growth? To our best knowledge, this is the first study dealing with the debt-growth linkage in CEE countries using the ARDL-based error correction (EC) model, which represents a significant contribution in this research area.

The paper is structured as follows. The introductory part is followed by a literature survey, covering the most relevant empirical studies and their results. The third section presents the implemented methodology, while the fourth section includes the data description. The fifth section shows the results and discussion, while the final part is the conclusion.

1. Literature Review

The earliest empirical debt-growth studies conducted by Barro (1979) and Sargent and Wallace (1981) show that a high public debt negatively affects the economic system, requiring both fiscal and monetary adjustments. The negative impact of public debt on the economic system is attributed to a phenomenon called public debt overhang, introduced by Krugman (1988) when modelling the behaviour of public debt-laden economies. It is defined as "the presence of an existing, inherited public debt sufficiently large that creditors do not expect with confidence to be fully repaid". This theory was later confirmed by Deshpande (1997) in his analysis of 13 severely indebted countries, which shows that high levels of public debt have a negative impact on investment. In addition, Sachs (1989) was among the first scholars to assume the existence of a non--linear relationship between public debt and economic growth, whereby public debt limits or reduces economic growth after reaching a certain turning point or upper limit. On the other hand, Schlarek (2004) applied a system generalized method of moments (system GMM) to developed countries over a thirty-year period (1970–2002), finding no statistically significant relationship between public debt and economic growth.

Increasing academic interest in EU countries started after Reinhart and Rogoff (2010) published their influential study, finding a strong relation between public debt and economic crises worldwide. They conclude that a high level of public debt (over 90% of the GDP) was associated with lower or even negative growth rates. Minea and Parent (2012) confirmed these findings saying that the debt-to-GDP ratio above 90% has a negative impact on economic growth, with significantly weakened effects after it reaches 115% of the GDP. Additionally, Eberhardt and Presbitero (2015) analysed 118 developing, emerging and advanced economies, revealing a negative relationship between debt and long-term macroeconomic performance. The own calculations also conclude that the complexity of the linkage prevents us from determining a universal threshold at which an economic slowdown occurs, due to sample heterogeneity. Similarly, Mencinger *et al.* (2015) emphasized a non-linear relationship for both emerging and advanced countries and a lower threshold for developing than for developed countries. Specifically, it ranges between 90 and 94% of the GDP for developed countries, and between 44 and 45% of the GDP for developing countries, since they are much more vulnerable to external shocks that affect foreign direct investment (FDI).

The presence of a non-linear relationship at public debt levels above 90% of the GDP is also confirmed by Woo and Kumar (2015). They argue in favour of appropriate economic measures to reduce public debt levels right before the point when it endangers economic growth. However, unlike Reinhart and Rogoff (2010), these own calculations show that a 10% increase in the public debt to GDP ratio leads to a decrease in per capita GDP of 0.2 % in developing countries, while the effect in developed countries is around 0.15%. The specified 90-percent fiscal threshold that is harmful for economic growth is also pointed out by Herndon *et al.* (2014) and Égert (2015). However, Herndon *et al.* (2014) argue that the effects are country-specific, implying that the threshold cannot be taken as an argument for the implementation of saving policies. Similarly, Égert (2015) re-estimated data used by Reinhart and Rogoff (2010) in a non-linear fashion and showed that a negative debt-growth relation is hard to determine uniquely. Namely, his results are sensitive to the sample characteristics and assumptions regarding the minimum number of observations required in each non-linear regime.

Most studies about the eurozone find a non-linear relationship between public debt and economic performance, but the results also differ significantly in terms of the threshold. For example, Checherita *et al.* (2012) and Baum *et al.* (2013) find a strong non-linear relationship between public debt and economic growth. In addition, the short-term impact is positive, with a very high significance until the public debt-to-GDP ratio reaches 67%; after reaching 95%, the impact is negative. Additionally, Dreger and Reimers (2013) show that the probability of a public debt crisis is much higher in the eurozone than in non -eurozone countries. Similarly, Alfonso and Alves (2015) indicate that joining the eurozone results in a 0.5% decrease in economic growth of the EU member countries, as well as a non-linear debt-growth nexus in a sample of 14 EU member states with the new public threshold at a 75% share of the GDP. These findings are also confirmed by Mencinger *et al.* (2014), but the turning point is country-specific and approximately at 80–94% for the old EU countries and 53–54% for the new EU members. The relationship is additionally confirmed by Dincă and Dincă (2015), but the public debt effects become negative at the 50% debt-to-GDP ratio. The difference in the turning points between new and old EU member states is caused by the fact that developed and developing countries are at different stages of their economic maturity (Mencinger *et al.*, 2015). Finally, Gómez-Puig and

Sosvilla-Rivero (2018) reveal a positive impact of debt on economic activity in the short -run and a negative effect in the long run, since debt reduces private investment and brings in uncertainty of tax policy.

By contrast, some studies have not determined a mixed debt-growth relationship, but instead a unilateral, positive or negative causality. For example, Čeh-Časni *et al.* (2014) conclude that the fiscal situation in central and Eastern European countries requires fiscal consolidation measures combined with pro-growth economic policies. This policy mix usually stabilizes public debt and mitigates its potential negative impact on economic activity. However, using a sample of 8 selected emerging market economies in the period 1980–2012, Fincke and Greiner (2015) show that public debt stimulates economic growth, with no identified detrimental effects. A potential explanation is that the countries are still not in a steady state as developed countries, but they are in a transition process characterized by high economic growth rates and growing investment in infrastructure. Also, they still have not reached the point at which the effects of public debt are negative, since the average debt -to -GDP ratio in these countries is significantly below the debt level in developed countries.

2. Data and Methodology

The data set includes quarterly macroeconomic variables, covering the period 2006Q1–2018Q4 with a total of 832 observations arranged in a panel data fashion. The sample consists of the sixteen central and Eastern European (CEE) countries according to the EU classification scheme (See Appendix, Table A1). The data are obtained from different sources such as the IMF's World Economic Outlook Database, the World Development Indicators (WDI) and country-specific statistical institutes. Before we estimate the model, we check the nature of the relationship, precisely whether the relation between the independent variables and the GDP is non-linear. For this pre-estimation check, we use the RESET test, following the two-step panel estimation procedure (Enders, 2015):

Step1: Estimate the best-fitting linear model. Let be the residuals from the model and denote the fitted values by $\hat{y}_{it}$.

Step 2: Select a value of H (usually 3 or 4) and estimate the regression equation:

$$e_{it} = \delta z_{it} + \sum_{h=2}^{H} \alpha_h \hat{y}_{it}^h \qquad for\, H \geq 2 \tag{1}$$

where z_t is the vector that contains the variables included in the model estimated in Step 1.

Once we check potential nonlinearity, we proceed to estimate the ARDL + EC model adjusted for panel data. The popularity of this methodology comes from the fact that cointegration of non-stationary variables is equivalent to an errorcorrection (EC) process, and the ARDL model has a reparameterization in the EC form (Engle and Granger, 1987; Hassler and Wolters, 2006). The ARDL (*p, q, …, q*) model has the following generic formulation:

$$y_{i,t} = c_0 + c_1 + \sum_{j=1}^{p} \phi_j y_{i,t-j} + \sum_{j=0}^{q} \beta_i^{'} x_{i,t-j} + u_{i,t} \tag{2}$$

where t = max(p, q), …, T assuming for simplicity that the lag of the order q is the same for all the variables in the design matrix $\mathbf{X}_{i,t}$. According to Pesaran *et al.* (2001), the variables s ($y_{i,t}$, $x'_{i,t}$)' are allowed to be purely I(0), purely I(1), or cointegrated. A reparametrized version of the model in the conditional errorcorrection (EC) form is given as follows:

$$\Delta y_{i,t} = c_0 + c_1 - \alpha\left(y_{i,t-1} - \theta x_{i,t-1}\right) + \sum_{j=1}^{p-1} \psi_{yj} \Delta y_{i,t-j} + \omega' \Delta x_{i,t} \sum_{j=1}^{q-1} \psi_i^{'} \Delta x_{i,t-j} + u_{i,t} \tag{3}$$

with the speed-of-adjustment coefficient $\alpha = 1 - \sum_{j=1}^{p} \phi_j$ and the long-run coefficient $\theta = \frac{\sum_{j=0}^{q} \beta_j}{\alpha}$. Following Pesaran *et al.* (2001), we use the AIC to decide about the optimal number of lags for each variable, while the cointegration test is performed using the bound F-test. The decision rule is quite simple: in the case of no cointegration, we are supposed to specify and estimate a panel ARDL model; alternatively, cointegration requires us to estimate a panel errorcorrection model (ECM).

Empirically speaking, our dependent variable is the GDP growth rate (annual %), whereas the design matrix $\mathbf{X}_{i,t}$ consists of several macroeconomic indicators for the country i in period t. Specifically, we include in the design matrix the initial GDP to address the possible convergence effect (GDP_0). In addition, the design matrix includes the following variables: public debt (*DEBT*) representing the general government gross debt calculated as a percentage of GDP; domestic savings (*SAV*), calculated as the gross domestic product minus final consumption expenditure (% of GDP); financial development (*FinDev*) proxied by domestic credit to the private sector (% of GDP); gross fixed capital formation (*FixCap*) as the share of the GDP serving as a proxy for investment; trade openness (*TRADE*) calculated as the ratio of export to GDP; the real interest rate (%) – the lending interest rate adjusted for inflation as measured by the GDP deflator (*IntRate*); the real effective exchange rate index (*ExRate*; 2010 = 100) and population growth (POPUL) to control for the labour supply. To be more specific, the real interest rate is the nominal interest rate adjusted for inflation; thus, we follow the ex-post approach because expected inflation rates are not available for many countries in the sample, especially for the Western Balkans.

Furthermore, the real effective exchange rate index is derived according to the World Bank methodology: the nominal effective exchange rate is divided by a price deflator or an index of costs. Also, the nominal effective exchange rate index is the ratio (expressed on the base 2010 = 100) of an index of a currency's period-average exchange rate to a weighted geometric average of exchange rates for currencies of selected countries and the euro area (World Bank, 2021). The structure of the design matrix is selected based on comparable recent studies (*e.g.*, Barro, 2003; Čeh-Časni *et al.*, 2014; Mencinger *et al.*, 2014; Alfonso and Alves, 2015; Bökemeier and Greiner 2015; Dincă and Dincă, 2015; Ferreira, 2016).

Also, we exclude public deficit (% of GDP) as a control variable for a couple of reasons: (1) public debt reflects a country's fiscal position more accurately (public deficit might be greatly affected by some non-economic transitory shocks such as the flood in Serbia in 2011 or political instability in North Macedonia in 2016); (2) we aim to emphasize the long-term orientation of the study; (3) to make the results comparable to the above mentioned influential studies in the field.

3. Results and Discussion

The descriptive statistics (See Appendix, Table A2) show a significant difference in mean and variability (measured by the standard deviation) between the variables. It seems that public debt, trade openness, exchange rate index, financial development and saving rate are highly volatile compared to the other variables. In addition, all the variables have non-zero skewness, meaning that their distribution has either a longleft tail (debt, GDP, trade openness, savings, interest rate and exchange rate) or a long right tail (all other variables). Similarly, the kurtosis shows that some distributions are peaked (leptokurtic) relative to the normal distribution (fixed capital, GDP, savings and population), while other variables have a flat (platykurtic) distribution relative to the normal. It is also worth mentioning that Croatia, Albania and Montenegro have the highest public debt, whereas Estonia, the Czech Republic and Slovakia have the lowest public debt. Also, Hungary, Slovakia and Slovenia have the highest export-to-GDP ratio, while Romania and Croatia have the lowest export-to-GDP ratio.

Another important phenomenon related to the negative gross domestic saving rate drew our attention. It happened in Montenegro in the period 2006–2013 (−2.09% on average) and Bosnia and Herzegovina in the period 2006–2016 (−5.66% on average). These negative gross saving rates imply that the final consumption expenditures in these two countries exceeded the GDP in the corresponding periods. This phenomenon (aggregate dissaving) in both countries was driven by high credit availability and much higher government spending relative to total budget revenues. Abundance of borrowed money fuelled excessive domestic demand and consequent explosion in domestic consumption, which in turn created a significant current account deficit. Consequently, both countries experienced consumption-driven short-run growth; however, it proved to be short-sighted, since these trends caused unsustainable and high leverage (indebtedness), public debt and external deficit. All other countries have positive gross domestic savings (as % of GDP) over the observed period (Estonia 29.47%, Hungary 27.40%, and Slovenia 26.87% on average).

The nonlinearity test results presented in the Appendix (See Table A5) show that the regression in Step 2 does not have any explanatory power, since none of the coefficients are significant and the coefficient of determination is almost zero. Consequently, we are dealing with a truly linear relationship, although the RESET test has little power to detect asymmetric models, since it uses integer powers of the fitted values. Our results are contrary to the findings of some other studies that revealed a non-linear and concave relationship (Reinhart and Rogoff, 2010; Minea and Parent, 2012; Mencinger *et al.*, 2014; Alfonso and Alves, 2015; Dincă and Dincă, 2015; Woo and Kumar 2015). Thus, we do not find evidence of non-linearity, meaning that we can proceed with a linear ARDL pre-estimation. It does not mean that further research agenda on this topic should not include some non-linear estimations, primarily because there are many other comparable empirical studies that implement a version of the threshold model to discover a cut-off point that would give a landmark for understanding the relationship between public debt and economic growth. A standard pre-estimation ARDL procedure (unrestricted constant and no trend model) requires testing for cointegration using the F-bound test, and the results are presented in Table 1.

Table 1: F-bound test (Unrestricted constant and no trend)

Test Statistic	Value	Significance	*i*(0)	*I*(1)
		***Asymptotic: n* = 10,000**		
F-Stat	7.925631	10%	3.27	4.24
		5%	3.89	4.45
		2.5%	4.51	5.62
		1%	5.26	6.42
Null hypothesis	***No levels relationship (no cointegration)***			
Conclusion:	*Cointegration exists*			

Source: Own calculations

Table 1 presents the F-bound test results, challenging the null hypothesis of no cointegration, *i.e*., the coefficients of the error correction terms are not all zero. As we can see, the *F*-statistic exceeds critical values for all significance levels. Thus, we reject the null hypothesis, meaning that there is a cointegration between GDP growth and the independent variables that are included in the preliminary ARDL model. The results suggest that we can estimate the long-run ECM model to determine whether there is a statistically significant long-run effect of the independent variables on GDP growth. Empirically speaking, we are supposed to estimate a one-lag (it is determined as an optimal lag using the AIC) ARDL errorcorrection model that is specified as follows:

$$\Delta GDP_{i,t} = \beta_0 + \sum_{j=1}^{p}\beta_{1j}\Delta GDP_{i,t-j} + \sum_{j=1}^{q}\beta_{2j}GDP_{i,j} + \sum_{j=1}^{q}\beta_{3j}\Delta DEBT_{i,t-j} + \sum_{j=1}^{q}\beta_{4j}\Delta SAV_{i,t-j}$$
$$+ \sum_{j=1}^{q}\beta_{5j}\Delta FinDev_{i,t-j} + \sum_{j=1}^{q}\beta_{6j}\Delta FixCap_{i,t-j} + \sum_{j=1}^{q}\beta_{7j}\Delta TRADE_{i,t-j} \tag{4}$$
$$+ \sum_{j=1}^{q}\beta_{8j}\Delta IntRate_{i,t-j} + \sum_{j=1}^{q}\beta_{9j}\Delta ExRate_{i,t-j} + \sum_{j=1}^{q}\beta_{10j}\Delta POPUL_{i,t-j} + \lambda ECT_{t-1} + \varepsilon_{i,t}$$

where ECT_{t-1} is the error-correction term, $GDP_{i,t-1}$ is the lagged GDP growth rate, $GDP_{i,j}$ is initial GDP that captures convergence effects, whereas the optimal number of lags (*j*) ranges from 1 to *p* (for the lagged dependent variable) or *q* (for all other independent variables).

Since the long-run effects on GDP growth are obtained through the estimation of the level equations, one might question the results regarding a possible problem of non -stationary variables at level. Namely, if the variables are found to be non-stationary but cointegrated (meaning that a linear combination of non-stationary variables is stationary), then the estimation still provides plausible and consistent long-run coefficients.

Apart from the F-bound test that suggests cointegration, we run the Pedroni (Eagle-Granger based) residual cointegration test (both within-dimension and between -dimension) to challenge the null hypothesis of no cointegration. Based on the results (panel PP-stat = 0; group ADF-stat = −2.5985, p = 0.0047), we strongly reject the null hypothesis, meaning that there is cointegration among the variables. Additionally, multiple stationarity tests suggest that the variables are non-stationary at level (See Appendix, Table A5); therefore, there is a need to specify and estimate an ECM model to reconcile the long -run behaviour of the variables with their short-run responses. To sum up, multiple stationarity tests indicate non-stationarity at level (See Appendix, Table A5), but the cointegration property allows us to estimate a plausible model since deviations are transitory and dissipate in the long run with a certain speed of adjustments. Note well that all the variables are stationary at first difference (See Appendix, Table A5), implying that a first-difference model that aims to capture the short -run effects also seems to have plausible and consistent coefficient estimates.

Following the methodological framework of the ARDL ECM model, it is interesting to look initially at the estimated long-run causal effects of the independent variables on GDP growth (see Table 2).

Table 2: Long-run effects and causality testing results

Panel A: Long-run effects (levels eq.)		**Panel B: Dumitrescu-Hurlin causality test**	
Variable	**Coefficient (SE)**	**W-Stat/Zbar-Stat (Prob.)**	**Conclusion**
GDP(0)	−0.0263 (0.0167)	***H_0: Variable does not homogenously cause GDP***	
DEBT	−0.2177** (0.0994)	6.3264/3.2980 (0.0010)	*Reject H_0*
SAV	0.1110*** (0.0365)	5.7166/2.7277 (0.0064)	*Reject H_0*
FinDev	0.0222 (0.0172)	1.7542/-0.9780 (0.3281)	*Fail to Reject H_0*
FixCap	0.1471** (0.0561)	8.7054/6.3051 (0.0000)	*Reject H_0*
TRADE	0.1786*** (0.0465)	5.5567/2.5781 (0.0099)	*Reject H_0*
IntRate	−0.0120** (0.0034)	5.9856/2.9792 (0.0029)	*Reject H_0*
ExRate	−0.0207 (0.0061)	7.9655/4.8309(0)	*Reject H_0*
POPUL	0.0929*** (0.0344)	6.6055/3.5590 (0.004)	*Reject H_0*
Model diagnostics			
$Mean_{GDP}/S.D._{GDP}$	2.655/3.869	SE of Reg.	0.3821
***R*-sq (Adj. *R*-sq)**	0.1154 (0.1038)	***F*-Stat/Prob.**	9.1006/0.0035
Serial correlation (BG Test: *no serial correlation***)**		**Chi-sq:** 1.574	**Prob:** 0.2102
Heteroskedasticity (BP test: *constant variance***)**		**Chi-sq:** 0.752	**Prob:** 0.3858

Significance: 1% (***), 5% (**), 10% (*)

Source: Own calculations

Based on the model diagnostics, we conclude that there is no serial correlation among the residuals since the Breusch-Godfrey (BG) test fails to reject the null hypothesis. Also, the Breusch-Pagan (BP) test fails to reject the null hypothesis of constant variance among the residuals, meaning that there is no heteroskedasticity.

The results presented in Table 2 indicate that most of the independent variables affect economic growth in the long run, *i.e.*, the results confirm a long-run causal relationship. To start with, the initial GDP proved to be statistically insignificant in this specification; accordingly, we cannot derive any conclusions regarding convergence/divergence in growth rates among CEE countries. However, we do want to emphasize that many extensive empirical studies support conditional convergence among similar countries, but usually do not support absolute convergence when it comes to heterogenous research samples (see Barro and Sala-i-Martin, 1992; Kumar and Woo, 2010). Additionally, there is a negative long-run causal relationship running from public debt and interest rate to GDP growth. These results are not surprising since many studies suggest that higher public debt implies higher tax burden in the future and higher interest rate due to the crowding -out effect, which in turn decreases consumption, investment, and economic growth (See Reinhart and Rogoff, 2010; Salmon, 2021). Higher saving rates have a positive long-run effects on growth (this finding is also supported by Thornton, 2009 and Guma, 2016), meaning that accumulation of financial resources helps the economy to grow faster probably through greater capital investment as a productivity booster. Furthermore, financial development is not statistically significant for long-run economic growth, which is contrary to the findings of some comparable studies (*e.g.*, Greenwood and Jovanovic, 1990; Claessens and Laeven, 2005); however, we reveal that growth matters for financial development (it is formally proved on the following pages), which corresponds to many other influential studies that promote the growth-driven view (*e.g.*, Rajan and Zingales, 1998; Zang and Kim, 2007). Simply speaking, our results support the demand-following hypothesis at the expense of the supply-lending hypothesis.

Fixed capital formation and trade openness have a positive long-run impact on growth, implying that greater capital investment and export-oriented development strategy might be helpful for CEE economies to catch up with the advanced countries. Finally, population growth causes higher growth rates in the long run mainly through higher aggregate inputs in the aggregate production function and higher consumption. Exchange rate seems to have a significant and negative effect on growth, implying that higher real effective exchange rate contributes to lower trade competitiveness, which in turn has detrimental effects on growth. The latter finding is in line with many comparable studies supporting this transmission channel using different parametric model specifications (See Rodrik, 2008; Vieira *et al.*, 2013).

The estimated long-run effects are additionally challenged using the pairwise Dumitrescu-Hurlin panel causality tests against the null hypothesis that a variable does not homogenously cause GDP growth rate. In all the cases (except for financial development) we reject the null hypothesis at the 5% significance level, meaning that there is a causality running from the independent variables (excluding financial development) to the GDP growth rate. Accordingly, the pairwise Dumitrescu-Hurlin panel causality results strongly support the level-based model of the long-run effects. We also test possible inverse effects, *i.e.*, the causality coming from the GDP growth to the government debt, savings and financial development. We fail to reject the null hypothesis in the case of the government (*W-Stat* = 3.03184; *Zbar-Stat* = 0.21682; *Prob* = 0.8283), and saving (*W-Stat* = 2.2947; *Zbar*-S*tat* = −0.4726; *Prob* = 0.6365), whereas the null hypothesis is strongly rejected in the case of financial development (*W-Stat* = 21.833; *Zbar-Stat* = 17.8004; *Prob* = 0). Thus, there is no reverse causality in each case, and the latter results lend support to the demand-following hypothesis, assuming that higher growth leads to financial development. To make the results as comprehensive as possible, we proceed to estimating the first-difference ECM model that provides an insight into the short-run effects of selected variables on economic growth (see Table 3).

The initial long-term estimation is followed by the short-term ECM estimation, which includes typical marginal effects, coupled with the speed of adjustment that is represented by the EC term (the cointegrating equation coefficient). Namely, we confirm a long-run relationship between the independent variables and economic growth, with a significant speed of adjustment (around 34.37% of disequilibrium dissipates before the next period). The short-run marginal effects seem to be congruent with the estimated long-term effects indicated above. Namely, the lagged GDP growth rate has a significant effect on current economic growth in the shortrun, which implies that economic growth is a short-run self-fuelling phenomenon. Although this indication might imply some level of divergence in growth rates, it is not supported by the empirical results (cannot be rejected either) since the initial GDP seems to be statistically insignificant in this specification. Thus, we have to remain inconclusive regarding convergence/divergence in economic growth among CEE countries, leaving room for further empirical investigations.

Table 3: ECM results (EC form)

Variable	Coefficient (SE)
***ΔGDP*(−1)**	0.1515*** (0.0122)
***GDP*(0)**	−0.1098 (0.0839)
ΔDEBT	−0.0271*** (0.0112)
ΔSAV	0.1057*** (0.0262)
ΔFinDev	0.1678 (0.1189)
ΔFixCap	0.1266** (0.0559)
ΔTRADE	0.2357*** (0.0871)
ΔIntRate	−0.1583** (0.0649)
ΔExRate	**−0.0339**** **(0.0163)**
ΔPOPUL	0.1433** (0.0619)
***CointEq*(−1)**	−0.3437*** (0.1128)
***R*-sq (Adj. *R*-sq)**	0.1527 (0.1186)
SE of Reg.	0.1675
***F*-Stat/Prob.**	4.4823/0

Significance: 1% (***), 5% (**), 10% (*)
Source: Own calculations

The remaining variables have a significant impact on the GDP growth rate at the 5% significance level (except for financial development, which is not significant), with quite predictable quantitative effects. Specifically, a 1% increase in the debt-to-GDP ratio would decrease the GDP growth rate by 0.03%; also, a unit increase in the real interest rate would decrease the GDP growth rate by 0.16%. Thus, higher debt and interest rate lead to lower economic growth in the short run; simply speaking, extensive government borrowing bids up the interest rate (the costs of borrowing) reducing consumption, investment and aggregate demand. A negative debt-growth linkage is also found in some similar studies such as Čeh-Časni *et al.* (2014), Bökemeier and Greiner (2015) and Ferreira (2016). A possible strategy in this situation is to optimize government spending to keep public debt below a critical level (a pre-defined threshold primarily set by the Maastricht Treaty and the international financial institutions); also, it might be more realistic for CEE countries to decrease their costs of borrowing through financial sector reforms and further liberalization. The latter would contribute to greater competition among financial institutions and would also attract more savings to drag down the real interest rate.

There are many other variables that have a significant short-run positive effect on economic growth. Specifically, with a 1% increase in the saving rate, we would expect GDP growth to increase by 0.11%. Thus, accumulation of financial resources at the countrylevel would lead to higher growth rates through efficient financial intermediation, given that highly efficient financial intermediaries channel these resources into the most lucrative investment alternatives. This finding is also supported by many comparable studies dealing with the impact of financial development on economic growth (Levine *et al.*, 2000; Christopoulos and Tsionas, 2004; Beck and Demirguc-Kunt, 2006). Additionally, a 1% increase in fixed capital would cause GDP growth rate to increase by 0.13%, whereas financial development is proved to be statistically insignificant, confirming our assumption about the growth-driven standpoint. Thus, investment (proxied by fixed capital formation) obviously contributes to economic growth, meaning that governments in CEE countries should boost economic incentives (primarily tax incentives) and overall institutional settings (government efficiency, political stability, economic freedom, etc.) to encourage greater fixed capital investments. The results also favour trade openness since trading partners benefit due to different competitive advantages and opportunity costs. To be more specific, with a 1% increase in the export-to-GDP ratio (% of GDP), we would expect the GDP growth rate to increase by 0.24%. Thus, CEE countries with stronger international economic relationships and export-oriented sectors tend to grow faster compared to those oriented on domestic markets and local trade arrangements. The latter finding suggests that CEE countries might benefit from further trade liberalization and single market policy that encourages cross-country economic cooperation.

Furthermore, the short-run effects of the real effective exchange rate correspond to the long-run estimation. Namely, the real exchange rate appreciation would cause net exports to decrease, with obvious detrimental effects on economic growth: a unit increase in the real exchange rate index would cause the GDP growth rate to decline by 0.03%. Consequently, CEE countries might be interested in keeping their exchange rates slightly undervalued to make their exports more competitive on the global markets. Finally, a unit increase in population growth would cause the GDP growth rate to increase by 0.14%; thus, population growth contributes to economic growth through lower input costs and a significant increase in labour force as an input in the aggregate production function;

moreover, larger populations usually consume more, which in turn increases aggregate output through the consumption multiplier.

Since the paper covers the period of the global financial crisis (2007–2012), and we are in the middle of the COVID-19 economic crisis, it would be interesting to derive the potential effects of these exogenous shocks on economic growth in light of the estimated short-run marginal effects. The global financial crisis (2007–2008) followed by a financial meltdown caused large-scale budget deficits and increasing public debts. These trends led to some serious economic consequences such as economic instability, lower savings, investments and exports. Additionally, these effects were observed globally due to financial globalization and cross-county economic interconnectedness. Central and Eastern European countries were also hit by reduced FDI due to increasing global economic and financial risks.

The impact of the COVID-19 crisis on the global economy (median global GDP dropped by 3.9% from 2019 to 2020) and CEE countries is also very significant. To reduce the imminent recessionary gap, the governments in CEE countries followed mandatory and voluntary restrictions proposed by the World Health Organization (WHO). Initially aimed at mitigating the impact on the workforce and the economy, these measures have already caused significant economic implications such as economic uncertainty, increasing public debt, international trade imbalances, etc. To be quite specific, the crisis resulted in significant economic deterioration of the CEE economies in 2020 (compared to 2019): (1) significantly higher unemployment: 0.92% on average (Montenegro: 2.78%, Albania: 1.86%, Estonia: +2.35%, Lithuania: +2.23%); (2) a sharp decline in export (% GDP): −5.19% (Croatia: −8.72%, Bulgaria: −8.61%, Albania: −8.19%, Slovakia: −6.82%, Bosnia and Herzegovina: −5.95%); (3) increasing recessionary gap (annual GDP growth rate): −4.58% on average (Montenegro: −15.31%, Croatia: −8.09%, Czech Republic: −5.79%, North Macedonia: −5.21%, Hungary: −4.67%); (4) moderate decline in gross savings (% GDP): −1.74% on average (North Macedonia: −5.19%, Bosnia and Herzegovina: −3.44%, Albania: −3.43%, Latvia: −2.63%, Croatia: −2.49%). These figures indicate that certain CEE countries (mainly Western Balkans) are particularly vulnerable to the COVID-19 induced macroeconomic shock. We can witness that economic recovery is still weak, and that the ongoing military conflict in Eastern Europe might produce a large-scale supply shock (due to the ongoing energy crisis) followed by negative spillover effects on the CEE economies.

Conclusion

The paper aimed to investigate the nature and magnitude of the effects of public debt as the main variable of interest, and a set of other macroeconomic (control) variables on economic growth (proxied by the GDP growth rate) in Central and Eastern European (CEE) countries in the period 2006Q1−2018Q4. The dataset included quarterly macroeconomic indicators collected from the IMF and the World Bank online databases, as well as from country-specific statistical institutes. The working panel included 832 observations, and we estimated the long-run and short-run effects using an ARDL-based EC model.

The estimated long-term effects are congruent with the short-term marginal effects; however, the magnitude and direction of the impact are variable-specific. Namely, higher public debt and real interest rate have detrimental long-term effects on economic development. It implies that CEE countries should implement a policy of tight budget control to minimize the current budget deficits and future public debt increases; the latter might require public sector reforms, especially in Western Balkan countries. Also, it might be possible to reduce the costs of borrowing by improving financial sector efficiency and attracting non-banking economic agents to increase savings. The results also indicate that higher saving rates and investment in fixed capital, coupled with trade openness, undervalued real exchange rates and population growth would contribute to economic development in CEE countries, in both the short and long run. Accordingly, CEE countries should promote tight budget constraints to minimize the crowding--out effect on the market for loanable funds. Another option would be to encourage both FDI and domestic private investments by providing non-monetary incentives and technical support, as well as to help domestic companies to get better access to EU structural and investment funds. Finally, these countries should encourage the private sector to be more export-oriented and implement pro-population social policy to support self-fuelling long-term economic growth. However, the COVID-19 economic crisis has already created many economic obstacles and constraints (supply chain breakdown, lower productivity, higher energy and transportation costs, higher unemployment, lower GDP growth, lower saving rates, etc.), which might significantly hamper future economic growth in CEE countries.

The study can be extended using some of the dynamic panel data estimators or using FE or mixed -effects frequentist and Bayesian estimators. Additionally, the dataset could be supplemented by using regional dummies and selected institutional indicators that seem to have a significant impact on growth, especially in developing countries.

Appendix

Table A1: Research sample (CEE countries)

No.	Country	No.	Country
1	Estonia	**9**	Bulgaria
2	Latvia	**10**	Slovenia
3	Lithuania	**11**	Croatia
4	Poland	**12**	Albania
5	Czech Republic	**13**	Montenegro
6	Slovak Republic	**14**	Serbia
7	Hungary	**15**	North Macedonia
8	Romania	**16**	Bosnia and Herzegovina

Source: Authors' calculation

Table A2: Descriptive statistics

Indicator/ Var.	*DEBT*	*FIXCAP*	*GDP*	*FinDev***	*TRADE*	*POPULATION*	*SAVING**	*IntRate***	*ExRate***
Mean	44.53133	23.484365	2.654937	50.26075	58.13181	−0.02592	17.7641	4.261970	97.625
Median	40.4	22.742232	3.025595	50.24430	62.82	−0.03387	20.6423	4.368389	99.4
Maximum	72.8	39.215849	11.965203	81.84658	77.13	0.260383	34.8168	12.670410	104.7
Minimum	12.4	15.949971	−14.83860	24.75368	33.57	−0.29781	−2.9436	−6.443934	85.4
Std. Dev.	19.3517	4.5610914	3.869585	11.47941	14.16855	0.12803	11.1367	3.764339	6.80393
Skewness	−0.11315	1.1261976	−1.474384	0.18826	−0.524513	0.196792	−0.8973	−0.409252	−0.78459
Kurtosis	−0.96662	1.3503262	4.980644	−0.08425	−0.75555	2.782367	0.1233	−0.170170	−0.32640
Observa-tions	832	832	832	832	832	832	832	832	832

* Negative gross saving rate (as % of GDP) implies that final consumption expenditures exceeded GDP for a specific period; Montenegro and Bosnia and Herzegovina had negative gross saving rates within the observed period.

** ***FinDev***: Credit to private sector (as % of GDP); ***IntRate***: Real interest rate - the lending interest rate adjusted for inflation as measured by the GDP deflator; ***ExRate***: Real effective exchange rate index (2010 = 100).

Source: Authors' calculation

Table A4: Testing for non-linearity (two-step procedure)

Step 1 Regression		Step 2 Regression	
Variable	**Coefficient (S. E.)**	**Variable**	**Coefficient (S. E.)**
Intercept	0.4770*** (0.0343)	***Intercept***	1.8490 (1.0399)
DEBT	−0.0339*** (0.0085)	***DEBT***	−0.0100 (0.0199)
SAV	0.0226*** (0.0027)	***SAV***	0.0071 (0.0204)
FinDev	0.0183 (0.0225)	***FinDev***	0.0468 (0.0387)
FixCap	0.0496*** (0.0191)	***FixCap***	0.0100 (0.0059)
TRADE	0.0677** (0.0242)	***TRADE***	0.0295 (0.0313)
IntRate	−0.11874** (0.05870)	***IntRate***	−0.0077 (0.0048)
ExRate	−0.0298*** (0.0057)	***ExRate***	−0.0109 (0.1384)
POPUL	0.0593*** (0.0089)	***POPUL***	0.2474 (0.2436)
***R*-sq (Adj. *R*-sq)**	0.3114 (0.2923)	***R*-sq (Adj. ADF *R*-sq)**	0.0623 (0.0453)
S. E. of Regression	0.0214	**S. E. of Regression**	2.8531
***F*-Stat/Prob.**	9.3245/0	***F*-Stat/Prob.**	0.4571/0.2301

Significance: 1% (***), 5% (**), 10% (*).

Source: Authors’ calculation

Table A5: Stationarity testing (level)

LEVEL										
	GDP		***Debt***		***FixCap***		***TRADE***		***Saving***	
Test	**Statistic**	**Prob.****	**Statistic**	**Prob.****	**Statistic**	**Prob.****	**Statistic**	**Prob.****	**Statistic**	**Prob.****
Null: Unit root (assumes common unit root process)										
Levin, Lin & Chu	−3.8061	0.0001	0.5999	0.7257	0.05990	0.5239	−3.3667	0.0004	−2.5941	0.0047
Breitung *t*-stat	−0.7290	0.2330	2.6285	0.9957	5.96430	1.0000	−2.3986	0.0082	−1.1071	0.1341
Null: Unit root (assumes individual unit root process)										
Im, Pes. and Shin	−0.93483	0.1749	0.8976	0.8153	4.29851	1.0000	−1.3733	0.0848	−0.7883	0.2153
ADF – Fisher Chi-sq	−20.6960	0.0550	7.4785	0.8244	15.1746	0.9949	19.4092	0.0791	15.9991	0.1913
PP – Fisher Chi-sq	18.7963	0.0936	19.6787	0.0734	27.7098	0.6836	43.8157	0	24.5762	0.0170
Conclusion	**Non-stationary**		**Non-stationary**		**Non-stationary**		**Non-stationary**		**Non-stationary**	
LEVEL										
	FinDev		***IntRate***		***ExRate***		***POPUL***		**–**	
Test	**Statistic**	**Prob.****	**Statistic**	**Prob.****	**Statistic**	**Prob.****	**Statistic**	**Prob.****	–	–
Null: Unit root (assumes common unit root process)										
Levin, Lin & Chu	−1.3350	0.0909	−4.79663	0	0.3002	0.6180	−1.4797	0.0695	–	–
Breitung *t*-stat	−0.3546	0.3614	0.75113	0.7737	1.6285	0.8921	1.4761	0.9300	–	–
Null: Unit root (assumes individual unit root process)										
Im, Pes. and Shin	−1.8730	0.0305	−1.38297	0.0833	7.8371	1.0000	−1.3250	0.0926	–	–
ADF – Fisher Chi-sq	38.6793	0.0772	26.5553	0.1482	7.9426	1.0000	19.9326	0.0684	–	–
PP – Fisher Chi-sq	28.9130	0	49.4686	0.0003	24.1178	0.9778	17.6986	0.1252	–	–
Conclusion	**Non-stationary**		**Non-stationary**		**Non-stationary**		**Non-stationary**		–	

** Probabilities for Fisher tests are computed using an asymptotic Chi-square distribution. All other tests assume asymptotic normality.

Source: Authors' calculation

Table A6: Stationarity testing (first difference)

FIRST DIFFERENCE										
	GDP		*Debt*		*FixCap*		*TRADE*		*Saving*	
Test	**Statistic**	**Prob.****	**Statistic**	**Prob.****	**Statistic**	**Prob.****	**Statistic**	**Prob.****	**Statistic**	**Prob.****
Null: Unit root (assumes common unit root process)										
Levin, Lin & Chu	–2.79034	0.0026	–11.600	0	–8.83214	0	–6.5513	0	–10.591	0
Breitung *t*-stat	–9.02411	0	–2.3784	0.0087	2.32386	0.9899	–3.9927	0	–0.8375	0.2012
Null: Unit root (assumes individual unit root process)										
Im, Pes. and Shin	–2.41540	0.0079	–3.6732	0.0001	–3.63718	0.0001	–2.6482	0.0040	–5.0243	0
ADF – Fisher Chi-sq	25.3836	0.0131	43.4797	0	69.5548	0.0001	28.8723	0.0041	38.3069	0.0001
PP – Fisher Chi-sq	58.2021	0	89.5967	0	78.5368	0	63.7823	0	50.8248	0
Conclusion	**Stationary**		**Stationary**		**Stationary**		**Stationary**		**Stationary**	
FIRST DIFFERENCE										
	FinDev		*IntRate*		*ExRate*		*POPUL*		–	
Test	**Statistic**	**Prob.****	**Statistic**	**Prob.****	**Statistic**	**Prob.****	**Statistic**	**Prob.****	–	–
Null: Unit root (assumes common unit root process)										
Levin, Lin & Chu	–4.3946	0	–9.27076	0	–3.0765	0.0010	–7.4759	0	–	–
Breitung *t*-stat	–5.5426	0	–1.26620	0.1027	–8.6704	0	–11.0488	0	–	–
Null: Unit root (assumes individual unit root process)										
Im, Pes. and Shin	–7.0202	0	–4.00028	0	–3.5783	0.0002	–9.2377	0	–	–
ADF – Fisher Chi-sq	67.8200	0	50.1354	0	23.7491	0.0006	101.1530	0	–	–
PP – Fisher Chi-sq	55.4980	0	77.6901	0	18.3115	0.0011	184.9010	0	–	–
Conclusion	**Stationary**		**Stationary**		**Stationary**		**Stationary**		–	

** Probabilities for Fisher tests are computed using an asymptotic Chi-square distribution. All other tests assume asymptotic normality.

Source: Authors' calculation

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