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Fiscal Sustainability Hypothesis Test in Central and Eastern Europe: A Panel Data Perspective

Abstract

This paper assesses the fiscal sustainability hypothesis for 10 Central and Eastern European countries (CEEC) between 1997 and 2019. The study adopts very recent panel econometric techniques which accounts for issues of structural breaks and cross-sectional dependence in the data generating process to examine the coin- tegration between government revenue and expenditures. Preliminary results show that revenues and expenditures do not have a long-run relationship and hence a rejection of the sustainability hypothesis. As a next step, we discriminate between structural and cyclical components of revenues and expenditures in order to place emphasis on the structural component. We argue that the structural component of fiscal variables represents the actual long term behaviour of the policymaker. Further results indicate that structural revenues and expenditures have a long-run relationship however with a slope coefficient less than unity which implies sustain- ability in the weaker sense. At that point, expenditures exceed revenues and if this continues for a long time the government may find it difficult to market its debts in the long run. This result suggests that the fiscal authorities in CEEC must therefore do more by taking long term actions to counteract the rising fiscal deficit problems.

Keywords

Fiscal Sustainability | Cointegration | Government Revenue | Government Ex- penditure

JEL Codes

H0, H6, E6

1. Introduction

The recent financial crises and global economic downturn prompted governments' Cinterventions across the world by way of fiscal expansions in attempts to stimulate aggregate demand [1]. This has implications on fiscal policy since spending must be financed by public deficits (Greiner & Fincke, 2015). Rising public deficits and debts to unsustainable levels may have long-run implications for the government since holders of government debts (usually the private sector) could lose confidence in government bonds. Secondly, the government could also default on its debts if it reaches unsustainable levels. The need to finance public deficit imposes a constraint on fiscal policy since governments in dynamically efficient economies have borrowing limits and face a present value borrowing constraint.[2] The issue of fiscal sustainability has therefore received considerable attention both in theoretical and empirical discussions. The fiscal stance is said to be sustainable if the future total discounted primary surplus in present value terms is equal to current debt. In other words, future stream of primary surplus when discounted in present value terms should be sufficient to offset the current level of public debt. Violating the conditions of the Intertemporal Budget Constraint (IBC) implies that debt will soar to an unsustainable level at a faster rate than the growth rate of the economy.

Prior to the accession of the European Union (EU), governments in Central and Eastern European

1 This principle is based on the Keynesian concept which has dominated the political-economic principles lately. The Keynesian school of thoughts advocates for government intervention to stabilize market economies.

2 See Abel et al. (1989) or a detailed discussion regarding dynamic efficiency of an economy.

Countries (CEEC) had to institute extensive fiscal policy actions to adjust their budgets and transform structures of revenues and expenditures whilst implementing institutional frameworks for fiscal policy reforms (Gleich, 2003). The objective was to ensure that they meet the necessary fiscal criterion in terms of size of debts, deficits and other obligations as stipulated in the Maastricht Treaty (MT) and Stability and Growth Pact (SGP). Eight CEEC out of the ten countries that joined the EU from the so-called eastern enlargement scheme had lower debt to GDP (Gross Domestic Product) ratios below the 60% threshold required by the MT and SGP; hence, Hallett and Lewis (2007) speculated that these CEEC could follow an explosive debt path for years without necessarily violating the fiscal sustainability requirements. Sixteen years after joining the EU, it remains to be seen if indeed these countries have pursued sustainable fiscal policies.

Most pioneered literature on fiscal sustainability started by empirically testing the stationarity of government debt and deficits (Westerlund and Prohl, 2010) as a way of fulfilling the government budget constraint. Notable among them are Hamilton and Flavin (1986), Trehan and Walsh (1988), Kremers (1988), Wilcox and Walsh (1989) and Baglioni and Cherubini (1993). Later authors such as Hakkio and Rush (1991), Lui and Tanner (1995), Quintos (1995), Ahmed and Rogers (1995) and more recently Afonso (2005) and Westerlund and Prohl (2010), have all zoomed in to specifically consider a more flexible approach of the cointegration between government revenue and expenditures. This is to ascertain if indeed revenues and expenditures have a long-run relationship with a positive cointegration vector, which is a confirmation of the sustainability hypothesis.

Even though a vast stream of empirical studies on fiscal sustainability on the European continent has been undertaken, there exists only a limited number of studies in the context of CEEC (Boekemeier & Stoian, 2018). For instance, Krajewski, Mackiewicz and Szymańska (1993 2016) examined the public debt sustainability for 10 selected CEEC countries using panel stationarity, a cointegration technique and a fiscal reaction function for the period 1990–2012. Their results indicated that the fiscal stance of selected CEEC countries is jointly sustainable. Similarly, Llorca and Redzepagic (2008) assessed the sustainability of fiscal policy for eight CEEC countries using panel cointegration analysis and found out these countries pursued sustainable fiscal policies for period 1999–2006 using quarterly data. Boekemeier and Stoian (2018) also investigated debt sustainability in 10 CEEC countries using estimates of a fiscal reaction function in its cubic form over the period 1998–2015. Their results revealed that government debts were at sustainable levels and that governments had not reached fiscal fatigue thresholds. Even though the studies above employed panel sustainability test for CEEC, none incorporated the possibility of structural breaks and cross-sectional dependence in the panel data generating process[3].

The aim of this paper is to ascertain the fiscal sustainability of 10 CEEC for the period 1997–2019 by investigating the long-run relationship between revenues and expenditure using panel cointegration.[4] The study makes use of a panel data analysis in order to benefit from the rich dynamism of panels. The availability of large macroeconomic datasets, over a long period of time and for different economies, is a recipe for a shift in the mean or trend of the individual time series. This increases the probability of break occurrence in the data (Carrion-i-Silvestre, Barrio-Castro & Lopez-Bazo, 2005). In cointegration analysis, structural changes have the tendency of affecting the cointegration vector, which is in contrast to conventional wisdom considering the fact that cointegration is a long-run stable relationship (Westerlund & Edgerton, 2008). This also leads to wrong inferences and hence proves how important it is to account for structural changes in the data generating process (Bai & Perron, 1998; Carrion-i-Silvestre et al., 2005).

Furthermore, cross country macroeconomic and financial datasets are associated with cross-sectional dependence because of inter-country links and dependencies, which is more of the rule now than an exception (Westerlund & Edgerton, 2008). Cross-sectional dependence affects the size properties of the unit root test and hence renders inferences incredible.[5] Hence, this study adopts the so-called 'second

3 The accession of some countries to the EU or Eurozone could represent a structural change in policy due to requirements that must be met and maintained by members of the union, notably requirements enshrined in the so-called MT and SGP.

4 These countries are Czechia, Estonia, Hungary, Latvia, Lithuania, Poland, Slovakia, Slovenia, Bulgaria and Romania.

5 Banerjee et al (2004) argued that unit root test which assumes cross-sectional independence suffers from size distortions as the actual size of the test is lower than the empirical size.

generational' econometric procedure which accounts for both cross-sectional dependence and structural breaks simultaneously in the data generating process, unlike other previous panel studies such as Beqiraj, Fedeli and Forte (2018), Claeys (2007) and Llorca and Redzepagic (2008).

Preliminary results show that revenues and expenditures do not have a long-term relationship and hence indicate a rejection of the sustainability hypothesis. Further, we discriminate between structural and cyclical components of revenues and expenditures in order to place emphasis on the structural component. This is the novelty of this paper when contrasted with previous panel cointegration sustainability studies between revenues and expenditures, such as Westerlund and Prohl (2010), Afonso (2005), Quintos (1995), Prohl and Schneider (2006), Claeys (2007) and Llorca and Redzepagic (2008). With motivations from Galí et al. (2003) who posited that the component of fiscal variables whose variations do not emanate from the influence of cycles represents discretionary fiscal policy, we use fiscal variables adjusted for cyclicallity.. As opined by Blanchard (2006), structural fiscal variables provide a benchmark by which fiscal policy can be judged. We argue that this structural component of fiscal variables (cyclically adjusted variables) represents the actual long-term behaviour of the policymaker and should be examined when conducting sustainability analysis. Further results indicate that the cyclically adjusted revenue and expenditure have a long run relationship. However, the slope coefficient of the cointegration relationship is less than unity and not strong enough to infer sustainability in the strong sense for cyclically adjusted variables[6]. These results suggest that even though cointegration exists for cyclically adjusted variables, the magnitude of the cointegration slope implies that expenditures are rising faster than revenue, which indicates fiscal deficits. Hence debt to GDP ratio is not bounded and if this continues for a long time, the debt stock will no longer be finite or sustainable, implying a weaker form of sustainability. There is therefore the need for the fiscal authorities in the selected countries to pursue long term actions that counteract rising fiscal deficits by way of fiscal consolidation to ensure the satisfaction of the government IBC. The contribution of the paper is in three folds. Firstly, it employs recent advances in panel econometrics that models structural breaks and cross-sectional dependence simultaneously in the data generating process for the sustainability hypothesis test in CEEC. Secondly, the study makes a case for the use of structural fiscal variables which is devoid of automatic response variables in the cointegration analysis for the sustainability hypothesis test. Finally, the study adds to the growing literature on fiscal sustainability in CEEC region.

The rest of the study is structured as follows. Section 2 will discuss the methodology used for the paper by laying the theoretical foundations for the sustainability test. The section will further discuss the various econometric procedures used for the test. Section 3 will provide the empirical estimation and discussion of the results. Section 4 concludes the paper.

2. Methodology

We begin with the government budget constraint which is assumed to hold at all times. A period government budget constraint in nominal terms is written as

$$g_t + i_t b_t - 1 = r_t + b_t - b_t - 1 \tag{1}$$

where g_t represents government spending, b_t is the government bond, i_t is the interest rates on bonds and r_t represents the government revenue. The above equation shows that the government expenditure (LHS) must be equal to total government receipts (RHS) at all times in order for the budget constraint to hold intertemporarily. Here we rule out the possibility of monetising of spending or the activities of monetary authorities. That is, we do not consider government printing money (also known as seignorage) to fund its expenditure as this is known to cause inflation. This assumption is plausible because the characteristic of modern economies is such that central banks independently control monetary policy (Greiner & Fincke, 2015) with little or no influence from fiscal authorities.

Taking the state of the economy into consideration and assuming i_t to be stationary around its mean i, Eq. (1) can be re-written as

$$\frac{g_t}{y_t} + \frac{(1+i)b_{t-1}}{y_t} = \frac{r}{y_t} + \frac{b_t}{y_t} \tag{2}$$

6 To infer strong sustainability in the sense of Quintos (1995), the cointegration slope must be equal to or greater than unity.

where y_t represents national income or nominal GDP. Simplifying further leads to

$$\frac{g_t}{y_t} + \frac{(1+i)b_{t-1}}{y_t}.\frac{y_{t-1}}{y_{t-1}} = \frac{r_t}{y_t} + \frac{b_t}{y_t} \tag{3}$$

$$\frac{g_t}{y_t} + \frac{(1+i)}{(1+f)}.\frac{b_{t-1}}{y_{t-1}} = \frac{r_t}{y_t} + \frac{b_t}{y_t} \tag{4}$$

where f is the nominal growth rate of the economy (GDP). Using capital notations, Eq. (4) can be rewritten as

$$G_t + (1 + \rho)B_{t-1} = R_t + B_t \tag{5}$$

where
$\frac{g_t}{y_t} = G_t,\ \frac{b_{t-1}}{y_{t-1}} = B_{t-1},\ \frac{r_t}{y_t} = R_t,\ \frac{b_t}{y_t} = B_t.\ \rho = \frac{i-f}{1+f}$
is the growth adjusted interest rate, which is assumed to be stationary for sake of simplicity.

Since further modification is required for empirical estimation, let $G^r_t = G_t + (\rho_t - \rho)B_{t-1}$, where ρ is the mean real interest rate and stationary.

Assuming that Eq. (5) holds continuously, then by forward substitution the present value budget constraint can be written as

$$B_{t-1} = \sum_{s=t+1}^{\infty}\left(\frac{1}{1+\rho}\right)^{s-t}(R_{s-1} - G^r_{s-1}) + lim_{s\to\infty}\left(\frac{1}{1+\rho}\right)^{s-t} B_{s-1} \tag{6}$$

Sustainability implies that the second term on the RHS of Eq. (6) converges to zero as time approaches infinity. This is also known as the transversality condition, which constraints the debt ratio not to grow at a faster rate than the interest rate[7]. If this is the case, then the current stock of debt should be equal to a total of both current and future discounted primary surpluses. As pointed out by Afonso (2005), the absence of no Ponzi condition can be tested empirically by testing the stock of debt for stationarity. Earlier studies that focused on testing the stationarity of public debt include Kremers (1988), Wilcox and Walsh (1989), Trehan and Walsh (1988) and Greiner and Semmler (1999).

Additionally, sustainability can be examined by testing the cointegration between revenues and expenditures, an idea initially pioneered by Hakkio and Rush (1991) and later Quintos (1995). Mathematically, this can be shown from Eq. (5), by making use of the auxiliary definition $GG_{t-1} = G_t + \rho B_{t-1}$. Assuming stationary real interest rate and applying the difference operator, the present value budget constraint can be re-written as

$$GG_{t-1} - R_{t-1} = \sum_{s=t+1}^{\infty}\left(\frac{1}{1+\rho}\right)^{s-t}(\Delta R_{s-1} - \\ - \Delta G^r_{s-1}) + lim_{s\to\infty}\left(\frac{1}{1+\rho}\right)^{s-t} \Delta B_{s-1} \tag{7}$$

Testing for the sustainability hypothesis can be done in two ways. First one could test for the absence of the no Ponzi scheme which implies that the second term of Eq. (7) approaches zero as time approaches infinity. Alternatively, we could assume the absence of no Ponzi scheme and test Eq. (7) directly. In this paper, we proceed to test the absence of no Ponzi scheme.

$$lim_{s\to\infty}\left(\frac{1}{1+\rho}\right)^{s-t} \Delta B_{s-1} = 0 \tag{8}$$

For Eq. (8) to hold, one-period government debt (ΔB_{s-1}) must not grow faster than the interest rate on debts. In order words, it is easier to see that Eq. (8) holds if ΔB_{s-1} is stationary as compared to a situation where it is not stationary. Considering that one period debt is given by the relationship $\Delta B_t = GG_t - R_t$, testing for stationarity of a one-period government debt implies testing for the difference stationarity for GG_t and R_t. This can be problematic if government spending and revenue are not stationary at their levels. However, if one can prove that they are stationary at their first difference, then the concept of cointegration can be applied. The intuition is that, if one variable can be written as a linear combination of the other with a slope coefficient such that the residual is proved to be stationary, then their relationship is stable and mean-reverting. In order words, the difference of these variables does not drift wide apart. Hence, we say they are cointegrated because they have a long-run stable relationship. From Eq. (8), this implies testing if GG_{t-l} and R_{t-1} are integrated of order 1 (I(1)) with an

7 Also known as the no Ponzi scheme, we rule out the possibility of the government issuing new debts in order to fund principal repayment and interest on existing debts.

imposition of cointegration vector (1, –1) as argued by Quintos (1995). One can test for cointegration equivalently as below:

$$R_t = \alpha + \gamma GG_t + \mu_t \tag{9}$$

Alternatively, making use of the expression $\Delta B_t = GG_t - R_t$, then from Eq. (9) we have

$$\Delta B_t = (1 - \gamma) GG_t - \alpha - \mu_t \tag{10}$$

Furthermore, $\gamma = 1$ implies sustainability since from Eq. (10) we infer that debt to GDP ratio is bounded and will grow at a constant rate. However, this condition was relaxed by Hakkio and Rush (1991), who demonstrated that the condition $0 < \gamma \leq 1$ guarantees sustainability if variables are cointegrated. Quintos (1995) argued further that $0 < \gamma \leq 1$ is both a necessary and a sufficient condition. She stressed that cointegration is only a sufficient condition for the sustainability hypothesis to hold.

At this point, it is important to make special remarks about the condition $0 < \gamma < 1$. Even though this is enough for sustainability, at this point the government expenditure exceeds its revenue and therefore the probability of default is high. It will be difficult to market its bonds and the government may have to pay high interest rates to issue new debt or attract new investors. Scenario $\gamma > 0$ guarantees sustainability since at this point, revenues are growing at a faster pace as compared to expenditures. Conversely at $\gamma < 0$, expenditures and revenues are moving in opposite directions and hence sustainability hypothesis is rejected. As shown by Quintos (1995), $\gamma = 1$ implies strong sustainability whereas $\gamma < 1$ implies some weaker form of sustainability. Therefore, the magnitude and sign of γ plays a major role in determining if, indeed, the sustainability hypothesis holds and the strength of the hypothesis.

Empirical cointegration test for Eq. (9) can be conducted conventionally by regressing R_t on GG_t simply by ordinary least square estimator (OLS) and testing the residuals for stationarity to confirm if cointegration holds. Westerlund and Prohl (2010) argued that such conventional test fails to reject the null hypothesis of no cointegration very often, which implies a rejection of sustainability hypothesis. They cited the problem of low power of the cointegration test because of low sample size. Panel datasets circumvent the power problem as it gives an opportunity to increase the sample size. Firstly, panels present more informative data because it has long sample size, provides more variability, involves less collinearity among variables and gives more degree of freedom for the model (Baltagi, 2008). Secondly, panel data affords researchers the opportunity to construct and test more advanced and complicated models as compared to time series or cross-sectional data, and finally panel helps to control for the effects of omitted variables bias in econometric (Hsiao, 2003). Hence the study will resort to a panel test which will subject the residuals in Eq. (9) to a cointegration test. The test is dynamic enough to account for structural breaks and cross-sectional dependence which is common to panel data analysis. The forthcoming sections will provide discussions on the econometric procedures for the panel test.

3. Empirical estimations and results

Firstly, we present a review of past empirical papers on fiscal sustainability with focus specifically on panel datasets. Subsequently, we will discuss our datasets and some characteristics of the data, after which we shall proceed with the empirical test of the fiscal sustainability hypothesis. Table 1 shows previous papers on panel data fiscal sustainability for mostly CEEC, Organisation for Economic Corporation and Development (OECD), and EU countries. Regarding CEEC, previous studies, notably by Llorca and Redzepagic (2008), Krajewski et al. (1993 2016) and Boekemeier and Stoian (2018), all point in the direction of a sustainable fiscal policy. It will therefore be interesting to compare our results directly with these studies.

Regarding our dataset, revenue, expenditure, and debt variables, these were all obtained from the OECD website for 10 CEEC[8]. All data exists in annual frequency. The sample period is from 1995 to 2019 and chosen based on the availability of data. A total of 250 observations are generated from a combination of 10 countries over a 25-year period. It is important to mention that we consider total expenditures, total revenues, and total debts as ratios of GDP.

8 These countries were chosen based on the availability of quality datasets and length of time series.

Table 1. Summary of existing empirical panel fiscal sustainability test

Reference	Sustainability test	Period and country	Findings
Afonso and Rault (2010)	Stationarity of debt and cointegration between revenue and expenditure	15 Selected EU countries (1970–2006)	Fiscal stance sustainability confirmed
Baldi and Staehr (2015)	Estimated fiscal reaction function of primary balance, debt and business cycle viables	Different groups of EU countries (2001–2004)	Sustainable fiscal stance for all groups post financial crises
Beqiraj et al. (2018)	Panel cointegration test between primary balance and public debt	21 OECD countries (1991–2015)	Fiscal stance judged to be unsustainable
Boekemeier and Stoian (2018)	Fiscal reaction function of primary balance and debt	CEEC (1997–2013)	Fiscal stance sustainable for selected countries
Brady and Magazzino (2018)	Stationarity of public debt	19 European countries (1970–2016)	Fiscal stance sustainability confirmed
Checherita-Wesphal and Žďárek (2017)	Fiscal reaction function of primary balance response to debt	18 Euro Area countries (1970–2013)	Sustainable fiscal stance
Claeys (2007)	Cointegration between revenue, spending and net interest payment	Selected European countries (1970–2001)	Sustainable fiscal policy
Krajewski et al. (2016)	Cointegration between revenue and expenditure and a fiscal reaction function	CEEC (1990–2012)	Sustainable fiscal stance
Lee et al. (2018)	Fiscal reaction function of primary balance response to debt	EU regional groups (1950–2014)	Varied results depending on the region
Llorca and Redzepagic (2008)	Cointegration between revenue and expenditure	CEEC (1999:1–2006:1)	Fiscal stance sustainable in selected coun tries
Prohl and Schneider (2006)	Cointegration between budget deficit and public debt	15 EU countries (1970–2004)	Fiscal stance sustainability confirmed
Westerlund and Prohl (2010)	Cointegration between revenue and expenditure	8 rich OECD countries (1977:1–2006:4)	Sustainability hypothesis confirmed for selected countries

CEEC, Central and Eastern European Countries; EU, European Union; OECD, Organisation for Economic Corporation and Development.

Figures 1 and 2 provide a graphical overview of revenue and expenditures as well as government debt for each of the countries in the panel. We noticed that in almost all of the cases, revenues and expenditure move in the same direction even though expenditures seem to be higher than revenue for most of the time periods. Poland, Hungary and Romania displayed high variability in the revenue-expenditure relationship. Moreover, the debt to GDP ratios of Hungary and Poland for most of the years exceed revenue and expenditures. For almost all the countries, we notice a rising public debt after 2008 which can be attributed to the activeness of fiscal policy within and after the financial crises. One can infer that since spending exceeded revenue, governments borrowed more to fund their increased spending. Figure 3 shows a scatter plot that reveals the relationship between revenues and expenditure with a smooth trend line. A positive upward-sloping relationship can be observed between the two variables, which provide some hints as to the nature of the relationship between the two fiscal variables.

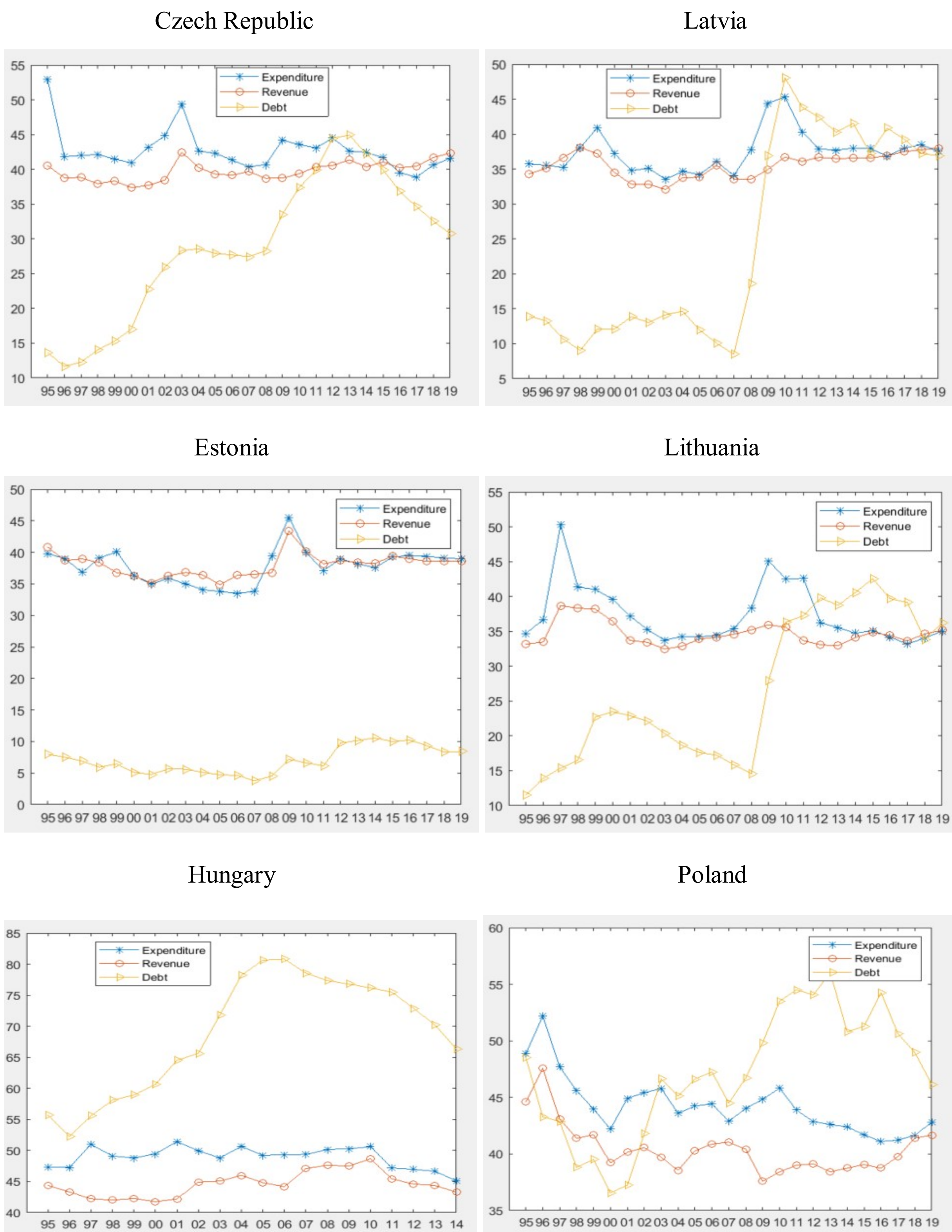

Figure 1. Revenue, Expenditure, and public debt

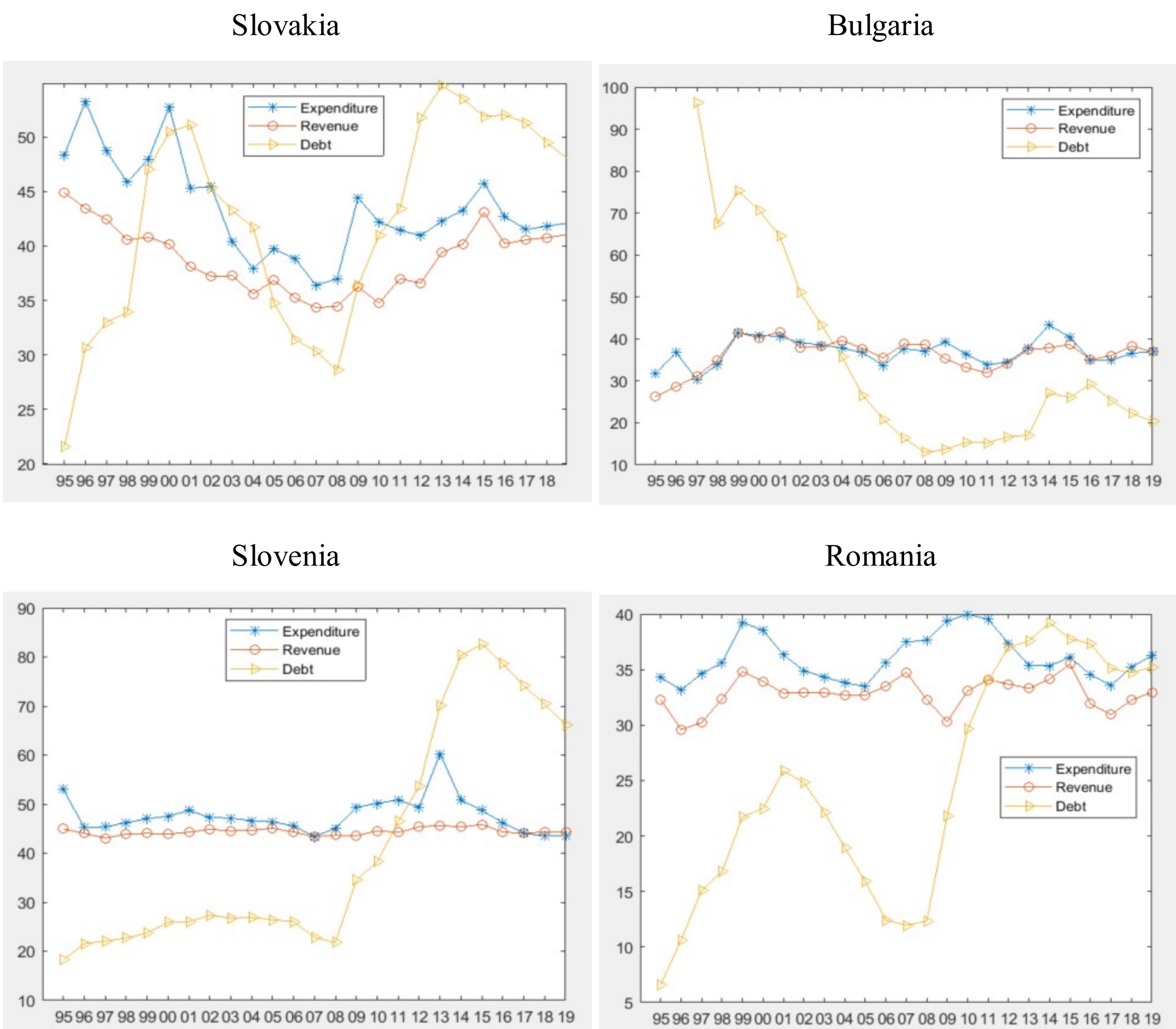

Figure 2. Revenue, Expenditure, and public debt

Figures 1 and 2 also provide some hints about the possibilities of structural breaks in the individual time series. Hence, it is feasible to test for the presence of structural breaks in the data. The presence of structural breaks could render statistical inferences erroneous if not accounted for in the data generating process. For instance, standard unit root tests are likely to exhibit biases towards non-rejection of the null hypothesis, hence leading to a wrong conclusion about results of the test (Carrion-i-Silvestre, et al., 2005). The issue of structural break has therefore received considerable attention in both theoretical and empirical econometric literature; notable among them includes Andrews, Lee and Ploberger (1996), Andrews (1993) and Bai and Perron (1998) among others. Structural breaks in the mean of data and the changes in the coefficient of a linear regression coincide with political, historical, and economic events (Zeiles et al., 2003) and are therefore not usually a random phenomenon.

To test the availability of structural breaks in the individual series, this study adopts the approach by Zeileis et al (2003). There they combined the F-statistics test by Andrews (1993) and Andrews and Ploberger (1994) to test the possibility of structural breaks in regression and the technique by Bai and Perron (2003) to locate the break dates and optimal breaks in the individual series of the data [9]. Table 2 provides results of the break dates for both revenue and expenditures. Regarding revenues, the number

9 Procedure is implemented in R studios with the package" 'strucchange'.". We select the optimal number of breaks by choosing the number of breaks with the least sum of square residuals.

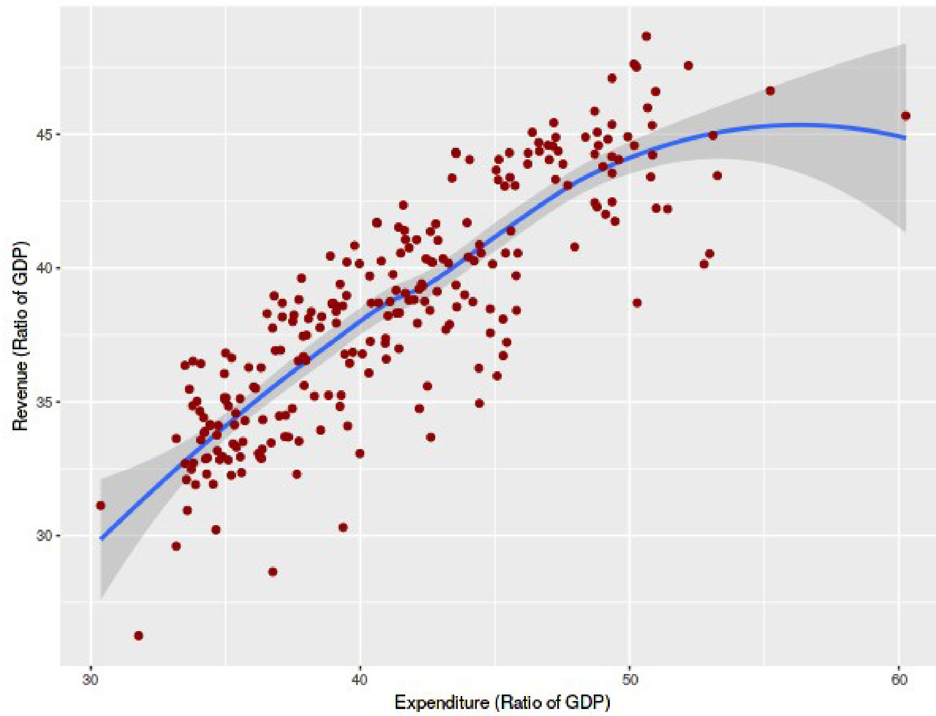

Figure 3. Panel scatter plot

Table 2. Dates of structural breaks for individual series

Countries (CEEC)	Revenue	Expenditure
Czechia	2002, 2010	2003
Estonia	1998, 2008, 2011	1999, 2007, 2010
Hungary	1997, 2006, 2011, 2015	1998, 2001, 2015
Latvia	1999, 2009	1997, 2000, 2008, 2011
Lithuania	2000	2000, 2008, 2011
Poland	1997, 2008, 2015	1997, 2011
Slovakia	1997, 2000, 2003, 2012	2002, 2008
Slovenia	2011, 2015	2008, 2015
Bulgaria	1997, 2008, 2012	1998, 2002
Romania	1997	1998, 2001, 2006, 2012

CEEC, Central and Eastern European Countries.

of breaks ranges between 1 and 4. We noticed that majority of the breaks were recorded before the early 2000s, which could possibly represent a policy shift as most of the CEEC were preparing to join the EU and therefore had to adjust their fiscal policies in order to meet the demands of the SPG and MT. Secondly, another break can be observed between 2007 and 2011 for most of the countries, which could also be attributed to the exogenous shock and the consequences from the global financial crises. This provides justification for the presence of the shocks and the fact that it must be accounted for in the data generating process.

Table 3 below provides a summary statistic of the panel dataset. We notice that there is more variability in expenditures as compared to the revenue components (from the standard deviation) over the sample period. Secondly, on the average, we observe that expenditures are higher than revenues, which is not so surprising since the role of government (spending) has become important especially in the 21st century either to stimulate economic growth or in direct response to macroeconomic shocks.

As per the SGP requirements, member states of the EU are required to maintain a strict upper limit of 3% deficit to GDP ratio (Wickens, 2008); hence, we investigate if member countries have followed this rule. Table 3 provides an overview of the deficit to GDP ratio of the CEEC during the sample period. We noticed that, with the exception of Estonia that violated the SGP only once (1999), all other countries violated this rule a couple of times. Firstly, it is observed that this occurred mostly between 1995 and 1998, which

Table 3. Panel summary Statistics

	Revenue-GDP ratio	Expenditure-GDP ratio	Debt-GDP ratio
Mean	0.391	0.419	0.340
Standard Deviation	0.039	0.505	0.178
Maximum	0.482	0.541	0.711
Minimum	0.308	0.321	0.038
Observations	250	250	250

is prior to their accession to the EU. Secondly, during the financial crises area between 2008 up till 2012, we also notice another round of SGP violation by all countries with the exception of Estonia. CEEC have therefore run fiscal deficits over the years and have not followed the 3% deficit limit rule strictly (Table 4).

Next, we test for evidence of cross-sectional dependence of the individual units in the panel. We deem it feasible to test for cross-sectional dependence, which is peculiar with macro panel data because countries in the same region respond to shocks in the similar ways, thereby generating serially correlated

Table 4. Deficit to GDP ratio

Year	Czh	Est	Hun	Lat	Lith	Pol	Svk	Slvn	Bulg	Rom
1995	**−12.44**	1.05	**−8.60**	−1.43	−1.53	**−4.26**	**−3.47**	**−8.15**	**−5.52**	−2.00
1996	**−3.01**	−0.32	**−4.38**	−0.42	**−3.22**	**−4.63**	**−9.82**	−1.09	**−8.11**	**−3.57**
1997	**−3.19**	2.15	**−5.55**	1.42	**−11.59**	**−4.61**	**−6.27**	−2.31	0.76	**−4.43**
1998	**−4.19**	−0.73	**−7.39**	0.03	**−3.03**	**−4.21**	**−5.30**	−2.33	1.08	**−3.24**
1999	**−3.14**	**−3.29**	**−5.23**	**−3.74**	−2.82	−2.28	**−7.17**	−2.97	0.09	**−4.42**
2000	**−3.57**	−0.04	−2.980	−2.73	**−3.17**	−2.98	**−12.63**	**−3.65**	−0.53	**−4.60**
2001	**−5.48**	0.20	**−3.94**	−1.95	**−3.51**	**−4.77**	**−7.22**	**−4.45**	1.05	**−3.46**
2002	**−6.36**	0.42	**−8.76**	−2.29	−1.85	**−4.85**	**−8.22**	−2.37	−1.16	−1.93
2003	**−6.89**	1.82	**−7.11**	−1.46	−1.27	**−6.08**	**−3.12**	−2.56	−0.39	−1.43
2004	−2.39	2.34	**−6.52**	−0.92	−1.39	**−5.04**	−2.32	−1.94	1.80	−1.09
2005	−2.9	1.08	**−7.72**	−0.36	−0.34	**−3.96**	−2.87	−1.32	1.00	−0.81
2006	−2.17	2.87	**−9.21**	−0.49	−0.27	**−3.56**	**−3.58**	−1.23	1.81	−2.14
2007	−0.65	2.73	**−5.03**	−0.51	−0.81	−1.85	−2.05	−0.05	1.10	−2.73
2008	−1.98	−2.65	**−3.73**	**−4.20**	**−3.09**	**−3.60**	−2.52	−1.39	1.59	**−5.35**
2009	**−5.45**	−2.16	**−4.69**	**−9.49**	**−9.13**	**−7.25**	**−8.15**	**−5.81**	**−4.05**	**−9.06**
2010	**−4.19**	0.19	**−4.39**	**−8.60**	**−6.92**	**−7.40**	**−7.46**	**−5.60**	**−3.13**	**−6.92**
2011	−2.73	1.06	**−5.19**	**−4.25**	**−8.95**	**−4.88**	**−4.46**	**−6.63**	−1.98	**−5.43**
2012	**−3.93**	−0.29	−2.27	−1.22	**−3.15**	**−3.74**	**−4.37**	**−4.0**	−0.32	**−3.65**
2013	−1.25	0.18	−2.54	−1.17	−2.61	**−4.18**	−2.87	**−14.58**	−0.43	−2.10
2014	−2.10	0.70	−2.76	−1.44	−0.62	**−3.65**	**−3.11**	**−5.51**	**−5.43**	−1.19
2015	−0.61	0.14	−1.97	−1.36	−0.27	−2.62	−2.67	−2.85	−1.72	−0.61
2016	0.72	−0.52	−1.76	0.06	0.23	−2.37	−2.48	−1.94	0.09	−2.62
2017	1.56	−0.77	−2.38	−0.52	0.45	−1.46	−0.95	−0.01	1.10	−2.64
2018	1.09	−0.56	−2.29	−0.74	0.60	−0.24	−1.06	0.77	1.75	−2.96
2019	0.75	−0.30	−1.83	0.51	0.13	−1.16	−1.03	0.72	−0.11	**−3.33**

Highlights in bold indicates violation of the EU SGP.

Source: Author's own computations.

EU, European Union; SGP, Stability and Growth Pact.

errors which affect inferences from the econometric test. Previous first generational econometric unit root test, such as Levin, Lin and Chu (2002), Im, Pesaran and Shin test (2003) and Maddala and Wu (1999); and cointegration test, notably Pedroni (2000, 2004) and Kao (1999), assume that the cross-section in the panel data is independent. Such The so-called first generational econometric test test usually suffers from size distortions, which affects the inferences (Banerjee, Marcellino & Osbat, 2004). Properly accounting for cross-sectional dependence in panels improves the efficiency of parameter estimates and simplifies statistical inferences (Hsiao, 2014). Two main cross-sectional dependence tests, namely Breusch–Pagan test and Pesearn test, are carried out in this paper. Proposed by Breusch and Pagan (1980), the test is based on a Lagrangian multiplier (LM), which is applicable to heterogeneous models and other variant panel models. Breusch–Pagan test is very convenient for datasets with short N and large T (Pesaran, 2004).

Table 5 presents results of the Breusch–Pagen and Pesaran cross-sectional dependence test. The null hypothesis for both tests indicates cross sectional independence in the panel datasets. In the case of expenditures to GDP ratio, there is a strong rejection of the null hypothesis for both tests. Hence, we accept the alternative hypothesis of cross-sectional dependence. For revenue to GDP ratio, there is a strong rejection (1%) for the Breusch–Pagan test and rejection at 10% significance level for the Pesaran test. Result provides evidence of cross-country dependence in the panel data and hence justifies the need to choose an econometric procedure that accounts for cross-sectional dependence.

From an econometric point of view, it is important to decide if, indeed, data can be pooled or not. According to Baltagi et al. (2008), imposing the pooling restriction reduces the variance of the pooled estimator. However, this could lead to a bias and hence wrong inferences if the restriction is false. Pooling data assumes that the parameters in the model are the same (homogeneous) across the individual countries. Similarly, we can also verify if the parameters are the same or different across the time periods. The decision of whether to pool or not is a natural question which arises in panel studies (Baltagi, 2005). Once the true nature of the parameters is established, it is then feasible to choose an appropriate econometric estimator. Considering an unrestricted model in the regression of the form

Table 5. Breusch–Pagan and Pesaran cross-sectional dependence test

Variables	Breusch–Pagan CD test	Pesaran CD test
	Chi-square, *P*-value	**z-value, *P*-value**
Expenditure to GDP ratio	107.84, 0.000	5.494, 0.000
Revenue to GDP ratio	113.42, 0.000	1.716, 0.0862

Pooled group variable by country. Null hypothesis of the test implies cross-sectional independence for both Breusch–Pagan and Pesaran test.
CD, cross-sectional dependence.

$$R_{it} = \alpha_i + \gamma_j GG_{it} + \epsilon_{it} \tag{11}$$

where R_{it} is government revenue to GDP ratio, G_{it} is government expenditure to GDP ratio, $_{it}$ represents the error term or residuals and α_i is the time-invariant intercept. From the slope coefficient, j could be the individual countries (heterogeneous across the countries) or time (heterogeneous across the time). In each case, we use the Chow (1960) test analogous to the F-test to test for poolability under the assumption that the residuals are normally distributed with a zero mean and constant variance. The test statistics is constructed by looking at the difference between the sum of squared residuals (SSR) of the restricted model and the SSR of the unrestricted model, and dividing by the SSR of the unrestricted model considering their degrees of freedom. In the context of panels, the unrestricted model could be a fixed effect (FE) within the model (with variable intercepts) or a pooled OLS regression model (constant intercept). A detailed discussion of the Chow test can be found in Wooldridge (2009) and Baltagi (2005).

From Table 6, if we consider an unrestricted model of a FE-within model with a time invariant slope, then clearly, we can reject the null of model stability or poolability at 1% significance level. Similarly, if we consider a pooled OLS regression with constant intercept and slope parameter, we can still reject the null hypothesis at 5% level. Hence, there is evidence to back the claim that we cannot pool the slope coefficient across the individual countries from our data. The slope coefficient is therefore heterogeneous across countries and cannot be considered as homogeneous. This provides useful guidance regarding the selection of an appropriate econometric model and procedure

Table 6. Chow test of poolability of coefficients (5% significance level)

Restricted model	Intercept	F-statistics	*P*-value	Verdict
FE (within)	Variable	2.8438	0.004	Not poolable
Pooled OLS	Fixed	14.892	0.000	Not poolable

Null hypothesis implies 'model stability' or constant coefficient.

FE, fixed effect; OLS, ordinary least square estimator.

suitable for accounting for heterogeneity in the slope coefficient.

As part of the cointegration requirement, the variables must be integrated of order 1. In other words, we test if revenues and expenditures are stationary at their first difference (I(1)). In this study, we adopt the Fourier unit root test by Nazlioglu and Karul (2017), which allows for smooth breaks in the mean of the series and cross-sectional dependence at the same time. This test is one of the few second generational unit root tests that accounts for both cross-sectional dependence and structural breaks. The test is a combination of an earlier test by Becker, Enders and Lee (2006) who employed a Fourier approximation function to model structural breaks and Hadri and Kurozumi (2011, 2012) who used a common factor structure to account for cross-sectional dependence. A Fourier approximation can be used to model structural shifts of any form or non-linearity in the deterministic term as this was shown by Becker et al (2006). It is important to note that the Fourier approximation is used to model breaks or shifts in a smooth gradual process which is in contrast to sharp breaks [10]. Another distinctive feature of the test is that the breaks are determined endogenously and do not have to be pre-determined.

The null hypothesis of the test implies 'stationarity' against the alternative of a unit root. The test depends on the Fourier frequency (k) which determines the swings and amplitude of the series. Considering the span of the time series in the panel (25 years), we choose $k = 3$, sufficient enough to cover the length of the time series. Tables 6 and 7 present the results of the panel univariate stationarity test and the test for the individual countries for expenditure and revenue, respectively[11]. We observe that the null of stationarity for the panel is strongly rejected at 1% significance level irrespective of whether we consider a model with a 'constant' or a model with 'constant and a trend'. For the individual countries, as k increases, we fail to reject the null of stationarity for most of the countries. However, considering the panel test statistic, the variables have a unit root and are therefore not stationary. It is necessary to test the first difference to ensure they are I(1). Stationarity test for the first difference of the variables shows the absence of unit root. Tables A1 and A2 in Appendix (for the sake of space) show the stationarity test results of the first difference of revenue and expenditures. It is observed that there is lack of evidence to reject the null hypothesis completely when we consider a model with a 'constant' and in some cases a 'constant and a trend'. As a robustness check, we employ the unit root test by Carrion-i-Silvestre et al. (2005), which accounts for structural breaks to test for if, indeed, the variables are I(1). Results in Table A3 in Appendix support the claim that revenue and expenditures are I(1). We therefore conclude that revenues and expenditure are stationary at their first difference.

After establishing that revenue and expenditure are I(1), we estimate the cointegration relationship between the variables. Regarding testing the cointegration relationship between revenues and expenditure, we resort to the test by Westerlund and Edgerton (2008). This test is very appealing because it serves as a one-stop-shop by accounting for structural breaks and cross-sectional dependence in panel data, making it very desirable. Secondly, the test is robust to serial correlation and heteroscedasticity in the residuals. Westerlund and Edgerton (2008) proposed two tests for the null hypothesis of no cointegration. The proposed test is derived from a LM function in the similitude of Schmidt and Phillips (1992), Ahn (1993), and Amsler and Lee (1995) unit root-based test.

We test the null of 'no cointegration' against an alternative hypothesis of 'cointegration' between revenue and expenditures. There are two proposed test statistics of the null hypothesis. The first test statistics $Z_\tau(N)$ is based on the least square estimate of the residual slope, whereas the second test statistics

10 Models with sharp breaks cannot be modelled by Fourier approximation. In such instances, dummy variables can be used to capture the sharp breaks. Carrion-i-Silvestre et al. (2005) developed a panel unit root test which is capable of accommodating sharp breaks in panels by using dummy variables.

11 I would like to thank Saban Nazlioglu for making the Gauss codes available.

Table 7. Panel stationarity test – Expenditure

Countries	Constant			Constant and trend		
	$k = 1$	$k = 2$	$k = 3$	$k = 1$	$k = 2$	$k = 3$
Czechia	0.070	0.309	0.302	0.050*	0.059	0.050
Estonia	0.099	0.410*	0.331	0.052*	0.124*	0.125*
Hungary.	0.203**	0.298	0.201	0.044	0.101	0.091
Latvia	0.183**	0.504**	0.414*	0.051*	0.080	0.083
Lithuania	0.052	0.055	0.116	0.051*	0.043	0.051
Poland	0.182**	0.496*	0.416*	0.053*	0.056	0.076
Slovakia	0.048	0.219	0.197	0.040	0.145**	0.139*
Slovenia	0.059	0.189	0.322	0.054*	0.097	0.085
Bulgaria	0.148**	0.089	0.095	0.051*	0.088	0.087
Romania	0.124	0.313	0.181	0.053*	0.136**	0.133*
Panel statistic	2.995***	3.508***	2.282***	4.938***	3.309***	2.460***
	(0.001)	(0.000)	(0.011)	(0.000)	(0.000)	(0.007)

Fourier panel stationarity test for 10 CEEC under the Null hypothesis of stationarity. *P*-values are for one sided test based on normal distribution.

Critical values (obtained from Becker et al. 2006, p. 289) for individual test statistics are as follows: 0.1318 (10%), 0.1720 (5%), 0.2699 (1%) for $k = 1$; 0.3150 (10%), 0.4152 (5%), 0.6671 (1%) for $k = 2$; 0.3393 (10%), 0.4480 (5%), 0.7182 (1%) for $k = 3$.

Critical values for constant and trend are as follows: 0.0471 (10%), 0.0546 (5%), 0.0716 (1%) for $k = 1$; 0.1034 (10%), 0.1321 (5%), 0.2022 (1%) for $k = 2$; 0.1141 (10%), 0.1423 (5%), 0.2103 (1%) for $k = 3$. Significance at 10%, 5% and 1% are denoted by *, ** and *** respectively.

CEEC, Central and Eastern European Countries.

$Z_\varphi(N)$ is based on estimating the t-ratio of the slope. A maximum of 3 breaks is chosen for the cointegration relationship. The selection of optimum lag length is based on an automatic procedure adopted from Campbell and Perron (1991) (Table 8).

Regarding the output of the test, we consider three scenarios. Firstly, we test the null hypothesis of 'no cointegration' under the condition of absence of breaks. That is, we assume there are no breaks in the cointegration relationship. Secondly, we test the null hypothesis by considering breaks in only the intercept (level break). Finally, we consider breaks in both the intercept and the slope (regime shift). From Table 9, we observe that none of the models are cointegrated when we consider significance at a strict 5% level. Considering a more relaxed significance level at 10%, we find evidence of cointegration for the model with no breaks for the $Z_\tau(N)$ test and no cointegration for $Z_\varphi(N)$. Hence, even with no breaks, the cointegration relationship is not strongly confirmed. This provides fresh evidence of lack of cointegration between total revenues and expenditures (all ratios of GDP)[12]. This implies a rejection of the fiscal sustainability hypothesis for CEEC, which is in contrast to previous studies on CEEC, notably by Krajewski et Al (1993 2016) and Llora and Redzepagic (2007). Even though they both employed a panel cointegration procedure, their studies did not to test for structural breaks and cross-sectional dependence in the cointegration relationships, which can be considered a major weakness. Hence, accounting for this dynamism (breaks and cross-sectional dependence) in a panel data setting reinforces the credibility of the results in this study.

3.1. Adjusting fiscal variables for cyclicality

Recall from Eq. (9) the cointegration relationship between revenues and expenditure. We decompose

12 I would like to thank Joakim Westerlund for making the Gauss codes available.

Table 8. Panel stationarity test – Revenue

Countries	Constant			Constant and trend		
	$k = 1$	$k = 2$	$k = 3$	$k = 1$	$k = 2$	$k = 3$
Czechia	0.146*	0.314	0.408*	0.047	0.048	0.068
Estonia	0.157*	0.271	0.108	0.044	0.066	0.091
Hungary.	0.342***	0.187	0.148	0.040	0.107*	0.103
Latvia	0.070	0.448**	0.313	0.052*	0.074	0.076
Lithuania	0.074	0.155	0.299	0.051*	0.076	0.054
Poland	0.170*	0.409*	0.352*	0.065**	0.101	0.103
Slovakia	0.262**	0.371*	0.362*	0.057**	0.153**	0.143**
Slovenia	0.056	0.128	0.239	0.056*	0.079	0.074
Bulgaria	0.268**	0.115	0.121	0.065**	0.114*	0.102
Romania	0.074	0.151	0.160	0.049*	0.121*	0.116*
Panel statistic	5.652***	2.716***	2.136***	5.606***	3.405***	2.523***
	(0.000)	(0.003)	(0.016)	(0.000)	(0.000)	(0.006)

Fourier panel stationarity test for 10 CEEC under the Null hypothesis of stationarity. *P*-values are for one sided test based on normal distribution.

Critical values (obtained from Becker et al. 2006, p. 289) for individual test statistics are as follows: 0.1318 (10%), 0.1720 (5%), 0.2699 (1%) for $k = 1$; 0.3150 (10%), 0.4152 (5%), 0.6671 (1%) for $k = 2$; 0.3393 (10%), 0.4480 (5%), 0.7182 (1%) for $k = 3$.

Critical values for constant and trend are as follows: 0.0471 (10%), 0.0546 (5%), 0.0716 (1%) for $k = 1$; 0.1034 (10%), 0.1321 (5%), 0.2022 (1%) for $k = 2$; 0.1141 (10%), 0.1423 (5%), 0.2103 (1%) for $k = 3$. Significance at 10%, 5% and 1% are denoted by *, ** and *** respectively.

CEEC, Central and Eastern European Countries.

Table 9. Panel cointegration test of revenue and expenditure – Europe

	$Z_\tau(N)$		$Z_\varphi(N)$	
Models	**Value (τ)**	***P*-value**	**Value (φ)**	***P*-value**
No breaks	−1.515	0.065	−0.963	0.168
Level break	0.879	0.810	0.690	0.755
Regime shift	0.422	0.663	0.437	0.669
Number of observations	250		250	

Westerlund and Edgerton (2008) cointegration test with three maximum number of breaks in the cointegration relationship, which are determined by grid search at the minimum of the SSR. Null hypothesis indicates 'No cointegration'. Displayed *P*-values are based on one-sided normal distribution test. *, ** and *** denote rejection of the null hypothesis at 10%, 5% and 1% respectively.

SSR, sum of squared residuals.

these fiscal variables into a trend and cyclical components. Following Galí et al. (2003), we posit that the cyclical component of fiscal variables necessitates automatic responses from government, which represents passive policy. In other words, this aspect does not constitute planned long-term government action and is influenced mainly by business cycles. The trend component on the other hand represents an active discretionary fiscal policy and hence should be only considered when examining the long-term behaviour of government policy. Decomposing Eq. (9), we have

$$R_t^c + R_t^\tau = \alpha + \gamma(GG_t^c + GG_t^\tau) + \mu_t \tag{12}$$

where R_t^c and R_t^τ represent cyclical and trend components of revenue whilst $GG^c_{\ t}$ and $GG^\tau_{\ t}$ are cyclical and trend components of government expenditures, respectively. Statistically, the cyclical component of variables is mean-reverting, and hence stationary. In other words, the cyclical component represents short-run dynamics, which will eventually die out in the long run. Secondly, in a panel cointegration set up, stationary and zero mean variables will end up in the residual term of Eq. (12) and will therefore not influence the cointegrating vector; hence, it is justifiable from an econometric perspective to exclude the cyclical component in the cointegration relationship (Beqiraj et al., 2018). Therefore, from Eq. (12), we end up with

$$R_t^{\tau} = \alpha + \gamma\left(GG^{\tau}{}_t\right) + \mu_t \tag{13}$$

A popular tool for decomposing a series into trend and cyclical component is the Hodrick Prescott (HP) filter (see Hodrick and Prescott (1997)). Consider a time series of the form

$$y_t = \tau_t + c \tag{14}$$

Using the HP filter, we denote mathematically by minimising the equation

$$Min_{\tau}\Big(\sum_{t=1}^{T}(y_t-\tau_t)^2+\lambda\sum_{t=1}^{T}\big[(\tau_{t+1}-\tau_t)-(\tau_t-\tau_{t-1})\big]^2\Big) \tag{15}$$

where y_t denotes the actual series at period t, τ_t denotes the trend component at time t and c represents the cyclical component of the series. Further, λ denotes the smoothing parameter, which is key to estimating the trend such that as λ approaches 0, the trend component approaches the actual time series, whereas λ approaching infinity implies a linear trend. Empirical value of $\lambda = 1{,}600$ was used by Hodrick and Prescott (1997) for US quarterly data. However for annual data, a value of $\lambda = 100$ is recommended (Martin, Hurn & Harris, 2013).

In a seminal paper, Hamilton (2018) proposed an alternative method for de-trending a series and proved that the HP filter is deficient in three respects. Firstly, he argued that HP filter imposes a spurious dynamic relationship which has no basis as far as the data generating process is concerned. Secondly, there are discrepancies between filtered values at the end of the sample and those at the middle of the sample and also spurious values. Finally, the values of HP smoothing parameter are vastly at odds with common practice and hence not reliable. To demonstrate his recommended approach, Hamilton (2018) applied OLS regression of series y_t on a constant and four recent values of y at time t as

$$y_{t+h} = \beta_0 + \beta_{1yt} + \beta_{2yt-1} + \beta_{3yt-2} + \beta_{4yt-3} + u_{t+h} \tag{16}$$

where y_t represents a quarterly time series and h is a eight quarter time horizon which is approximately 2 years [13]. The residual u_t which is assumed stationary represents the cyclical component of the original series y_t and is given by

$$u_{t+h} = y_{t+h} - \hat{\beta}_0 - \hat{\beta}_{1yt} - \hat{\beta}_{2yt-1} - \hat{\beta}_{3yt-2} - \hat{\beta}_{4yt-3} \tag{17}$$

The residual is stationary provided the fourth difference of y_t is stationary. The study will adopt both filters in order to ascertain if results after using filters differ significantly.

As a requirement for cointegration, we test all variables at their levels and first difference to ensure they are $I(1)$. Results in Tables A4–A7 in Appendix indicate that cyclically adjusted expenditure and revenue have a unit root in their levels and are stationary in first difference, paving the way for the cointegration test. Table 10 provides result of the cointegration test between cyclically adjusted revenues and cyclically adjusted expenditure. When we consider a model with no breaks, we rejected the null at a strict 1% level, which implies accepting the alternative hypothesis of cointegration. Secondly, there is strong cointegration when we consider breaks in the level or regime shift with a strong rejection of the null hypothesis of no cointegration at 1% level for both $Z_\tau(N)$ and $Z_\phi(N)$. The result is not any different if we use cyclically adjusted variables from the HP filter (see Table A8 in Appendix). The result strongly supports cointegration between cyclically adjusted revenue and expenditures for CEEC countries. This implies that if we consider cyclically adjusted variables, we can infer that governments in CEEC have jointly pursued a sustainable fiscal policy.

13 In the case of data with annual frequency, Hamilton (2018) recommended that the value of h should be fixed at h = 2

Table 10: Panel Cointegration cointegration test of cyclically adjusted revenue and cyclically adjusted spending

	$Z_\tau(N)$		$Z_\varphi(N)$	
Models	**Value (τ)**	***P*-value**	**Value (φ)**	***P*-value**
No breaks	−2.842***	0.002	−3.587***	0.000
Level break	−3.076***	0.001	−2.950***	0.002
Regime shift	−2.368***	0.009	−2.984***	0.001
Number of observations	220		220	

Westerlund and Edgerton (2008) cointegration test with three maximum number of breaks in the cointegration relationship which are determined by grid search at the minimum of the SSR. Null hypothesis indicates 'No cointegration'. Displayed *P*-values are based on one-sided normal distribution test. *, ** and *** denote rejection of the null hypothesis at 10%, 5% and 1% respectively. Hamiliton Filter is used to obtain cyclically adjusted revenues and expenditures.

SSR, sum of squared residuals.

3.2. Estimation of cointegration vector

Once cointegration is established, it is necessary to estimate the equilibrium parameter in the long run dynamic relationship in order to infer the sustainability hypothesis. Recall from Section 2 that to infer strong sustainability in the sense of Quintos (1995), the size of the slope coefficient must be equal to unity; otherwise a weak form of stationarity is inferred. It is important to note that when variables are expressed as ratios of GDP or in per capita terms, then it is even crucial to get a slope coefficient of 1 in the other to ensure that debt is bounded and does not explode to infinity (Afonso, 2005). We estimate a panel form of Eq. (13) with heterogeneous slope coefficient as below

$$R_{it}^{\tau} = \alpha_i + \gamma\left(GG^{\tau}{}_{it}\right) + \mu_{it} \tag{18}$$

Even though the OLS has been found to be super consistent, it has been proven to be deficient with finite data sets and complex dynamic relationships (Forest and Turner, 2013) and insufficient for panel data (Baltagi, 2005). Considering the fact that our slope coefficient in Eq. (18) is heterogeneous, it is necessary to use estimates that account for slope heterogeneity. This is quite common with macro datasets with large time series and cross-sectional dimensions represented by countries or regions. The asymptotic of macro panels are different from that of traditional micro panels with large number of cross-sections and a smaller time period. Such micro datasets are usually estimated with FE, random effects, instrumental variables techniques to account for possible endogeneity such as the generalised method of moments (Blackburne & Frank, 2007). However, these estimators rely on the assumption that the slope coefficient is homogeneous, something which is not applicable to our model.

Likely candidates are the mean group (MG) type of estimates, which includes the MG estimator, common correlated effect mean group (CCMG) and augmented mean group. The MG first developed by Pesaran and Smith (1995) is similar to the FE-within model; however, it averages the slope for each individual country in the panel. From Eq. (18), one would estimate the N-group specific ordinary least squares regression and average the estimated coefficient for the group to account for heterogeneity in the coefficient. From Eq. (18), MG estimator is as follows:

$$\hat{\gamma}_{mg} = \frac{1}{N}\sum_{n=1}^{N}\hat{\gamma}_{ols,n} \tag{19}$$

One of the main criticisms of the MG estimator is the fact that it does not account for the issue of cross-sectional dependence in panel data. Hence inferences from this estimator should be made with some caution due to potential bias. To circumvent this problem, Pesaran (2006) developed the CCEMG that accounts for cross-sectional dependence by allowing for heterogeneous impact across panel members. From Eq. (18), we expand the error term to include an unobserved common factor which is recovered by cross sectional averages of the dependent variable and independent variable:

$$\mu_{it} = \lambda F_{it} + \eta_{it}.$$

where F_{it} is the common factor term and η is a random shock. Once the unobserved common factor is recovered, the estimator of the group can be obtained by again averaging the slope coefficients across the panel in a similar fashion as the MG estimator. This estimator is therefore robust against cross-sectional dependence (see Pesaran (2006) for further details).

A further problem arises if the regressor in the model is potentially endogenous. Then most estimators which do not account for endogeneity bias will suffer from depending on nuisance parameter (Westerlund & Prohl, 2010). Authors such as Kao and Chiang (2000) and Chen, McCoskey and Kao (1999) therefore recommended the fully modified OLS (FMOLS) proposed by Phillips and Hansen (1990) and dyanmic OLS (DOLS) introduced by Saikkonen (1991) and later advanced by Stock and Watson (1993) as promising models for estimating the long-run vector in a cointegration regression.

Since our slope coefficient is very heterogeneous, MG estimator of FMOLS and DOLS will be appropriate in this context. Specifically, the MG-FMOLS and MG-DOLS introduced by Pedroni (2000, 2001) are suitable for estimating the cointegration vector such that it is consistent with cross-sectional heterogeneity in panel cointegration studies. Again, consider a panel regression of the form in Eq. (18)

$R_{it} = \alpha_i + GG'_{it}\gamma_i + u_{it}$ where R_{it} is the dependent variable (1×1) and γ is the vector of slope parameter, α_i represents the intercept, $i = 1,....,N$ and $t = 1,....,T$. u_t is the disturbance term which is assumed to be stationary. GG_{it} is a kx1 regressor vector which is assumed to follow the process,

$$GG_{it} = GG_{it-1} + \epsilon_{it}$$

Under the specification above, if R_{it} and GG_{it} are assumed to be cointegrated, they are both integrated process of order 1 (written for notational simplicity as $I(1)$). The OLS for Eq. (18) is given by;

$$\hat{\gamma}_{OLS} = [\sum_{i=1}^{N}\sum_{t=1}^{T}(GG_{it} - \bar{G}G_i)(GG_{it} - \bar{G}G_i)']^{-1}$$
$$[\sum_{i=1}^{N}\sum_{t=1}^{T}(GG_{it} - \bar{G}G_i)(R_{it} - \bar{R}_i)] \quad (20)$$

The FMOLS makes correction for the OLS model by accounting for endogeneity and serial correlation in the OLS in Eq. (18) by applying a non-parametric correction. The MG-FMOLS (accounting for heterogeneity in slope coefficient) is given by

$$\hat{\gamma}_{MG-FMOLS} = \Big[N^{-1}[\sum_{i=1}^{N}\sum_{t=1}^{T}(GG_{it} - \bar{G}G_i)(GG_{it} - \bar{G}G_i)']^{-1} * [\sum_{i=1}^{N}(\sum_{t=1}^{T}(GG_{it} - \bar{G}G_i)\hat{R}^{+}_{it} - T\hat{\Delta}^{+}_{\epsilon u})\Big] \quad (21)$$

where $\hat{R}^{+}_{it} = \alpha_i + GG'_{it}\gamma_i + u_{it} - \hat{\Omega}_{u\epsilon}\hat{\Omega}^{-1}_{\epsilon}\Delta GG_{it}$ is the endogeneity correction term, such that $\hat{\Omega}_{\epsilon}$ and $\hat{\Omega}_{u\epsilon}$ are consistent estimates of Ω and Ω_u, respectively, and where Ω is the covariance matrix of GG and R and $\hat{\Delta}^{+}_{\epsilon u}$ is the serial correlation correction term given by

$$\hat{\Delta}_{\epsilon u} - \hat{\Delta}_{\epsilon}\hat{\Omega}^{-1}_{\epsilon}\hat{\Omega}_{\epsilon u.}$$

The MG-DOLS regression on the other hand entails augmenting the cointegration model with lags and leads of ΔGG_{it}, so that there is orthogonality between the error term and the regressors. This corrects the endogeneity and serial correlation in the panel cointegration regression, and the concept of the MG is to account for the heterogeneity across cross sections. The MG-DOLS $\hat{\gamma}_{DOLS}$ is obtained from the equation

$$R_{it} = \alpha_i + GG'_{it}\gamma_i + \Sigma c_{ij}\Delta GG_{it+j} + \bar{\upsilon} \quad (22)$$

Where where $\bar{\upsilon}$ is the combination of the disturbance terms.

From Eq. (22), the panel DOLS is given as

$$\hat{\gamma}_{MG-DOLS} = \Big[N^{-1}\sum_{i=1}^{N}\Big(\sum_{t=1}^{T}z_{it}z'_{it}\Big)^{-1}\Big(\sum_{t=1}^{T}z_{it}\tilde{s}_{it}\Big)\Big] \quad (23)$$

where $z_{it} = GG_{it} - \bar{GG}_{it}, \Delta GG_{it-k},......,\Delta GG_{it+k}$ is a vector of regressors and $\tilde{s}_{it} = R_{it} - \bar{R}_i$.

The study makes use of all four models (MG, CCEMG, MG-FMOLS and MG-DOLS) to estimate the long-run cointegration vectors. Table 11 shows the estimated cointegration slope (γ) using the different estimators. In the case of MG-DOLS, we explore lags and leads from 1 to 4 to experiment with the sensitivity of the γ coefficient. For FMOLS, we make use of the Bartlett Kernel for the long-run covariance matrix. The long run coefficient for the MG and CCEMG are 0.499 and 0.364, respectively, and P-values indicate their statistical significance with low standard errors. When we consider the MG-FMOLS and MG-DOLS, the long-run coefficients are 0.938 and 0.935, respectively, which are higher comparatively. The probability values indicate that the estimated slope

Table 11. Long run coefficient for cyclically adjusted fiscal variables

Stat	MG	CCEMG	MG-FMOLS	MG-DOLS[a]
γ	0.499	0.364	0.938	0.935
Test stat	3.390	3.815	480.74	604.07
P-value	0.000	0.000	0.000	0.000
Std error	0.147	0.095	0.002	0.02
Obs	220	220	210	130
Shapiro–Wilk Normality test	0.970	0.977	0.992	0.986
P-value	(0.000)	(0.001)	(0.260)	(0.220)
Peseran CD test	3.652	−2.411	3.328	−0.260
P-value	(0.000)	(0.016)	(0.001)	(0.795)

[a]Reported lags and leads of 4 for MG-DOLS. The study explored lags and leads from 1 to 4; however, this does not change the estimates of the parameter.
CCEMG, common correlated effect mean group; CD, cross-sectional dependence; DOLS, dyanmic OLS; FMOLS, fully modified OLS; MG, mean group; OLS, ordinary least square estimator.

is statistically significant with low relative standard errors.

We conduct some residual diagnostic tests to ascertain address the question of which estimator performs better. Recall that the model Eq. (18) relies on the assumption that the residuals are normally distributed with a zero-mean and a constant variance. Hence it is feasible to test if, indeed, this is the case. The lower part of Table 11 depicts test statistics and probability values of the Shapiro-–Wilk test (see Royston (1982)). From the *P*-values, we can reject the null hypothesis of '"normality in residuals'" for the MG and CCEMG estimators at 1% significant level. In the case of MG-FMOLS and MG-DOLS, we cannot reject the null hypothesis of residual normality, ; hence, these two models perform better because their residuals are normally distributed.

Secondly, we conduct cross-sectional dependenking use of the Pesaran test (according to Pesaran (2004)). Reported *P*-values from Table 11 reveal that the null hypothesis of cross-sectional independence can be rejected for MG, CCEMG and MG-FMOLS at 1%, 5% and 1%, respectively, in favour of the alternative hypothesis of the presence of cross-sectional dependence in the residuals. In the cas of MG-DOLS model, we cannot reject the nul hypothesis even if we consider a lax 10% signiicance level. This provs that the MG-DOLS is robust against cross-sectional dependence and has normally distributed residuals making it the most efficient estimator among the tee.

Considering the size of the slope coefficient for the two efficient estimators (MG-DOLS and MG-FMOLS), it is important to establish if indeed they are equal to 1. Recall from Section 2 that a slope coefficient of 1 guarantees strong fiscal sustainability since it implies that the debt to GDP ratio is bounded. To ascertain if cointegration slope (γ) is indeed 1, it is plausible to conduct a hypothesis test of the coefficient. We employ the Wald test under the null hypothesis that $\gamma = 1$ ($H_{0:}\ \gamma = 1$), as against the alternative that $H_{a:}$ $\gamma < 1$. The Wald test statistics takes the form:

$$W = (\hat{\gamma} - \gamma_0)'[var(\hat{\gamma})]^{-1}(\hat{\gamma} - \gamma_0) \sim \chi_p^2 \qquad (24)$$

which reduces to

$$W = \frac{(\hat{\gamma} - \gamma_0)^2}{var(\hat{\gamma})} \sim \chi_p^2 \qquad (25)$$

where $\hat{\gamma}$ is the maximum likelihood estimate of the parameter to be tested, and γ_0 is the parameter which is assumed to be true under the null hypothesis. If the null hypothesis is true, then W is chi-square distributed with p degrees of freedom, which also represents the number of parameters to be estimated.

Results of the Wald test shown in Table 12 imply the rejection of the null hypothesis of a unit slope for the two models (MG-DOLS and MG-FMOLS) indicating that $0 < \gamma < 1$. This is statistically significant if we consider the *P*-values of the Wald test. Further, we construct confidence intervals to show the position of the true value of γ at 95% level. All evidence shows that $\gamma < 1$, which implies weak sustainability for cyclically adjusted variables in the sense of Quintos (1995). Even though cyclically adjusted revenue and expenditure are cointegrated, the magnitude of the cointegration slope is not strong enough to guarantee strong sustainability. The intuition is that considering a linear Eq. (18), an increase in expenditure by 1 unit will induce revenue to increase by less than 1 unit, all other things being equal. Hence expenditure to GDP ratio grows more than revenue to GDP, implying the accumulation of debts and hence a bubble debt term

Table 12. Wald test of coefficient and confidence intervals

Models	MG-FMOLS	MG-DOLS
T stat	−32.018***	−49.89***
Chi-square (1 df)	1,025.13	1,754.72
P-value	0.000	0.000
95% confidence intervals	(0.934–0.941)	(0.932–0.938)

*, ** and *** denotes rejection of the null at 10%, 5% and 1% respectively. Null hypothesis: $\gamma = 1$.
DOLS, dyanmic OLS; FMOLS, fully modified OLS; MG, mean group; OLS, ordinary least square estimator.

in the long run. Even though there is cointegration for cyclically adjusted variables, debt to GDP ratio is not finite in the long run. Hence, we refer to the fiscal stance of CEEC as weakly sustainable in the sense of Qunitos (1995). Based on the above findings, we therefore conclude that CEEC jointly have pursued a weakly sustainable fiscal policy if we consider cyclically adjusted revenues and spending to GDP ratio.

4. Conclusion

This study sought to ascertain if the fiscal sustainability hypothesis holds for 10 CEEC from the period 1995 to 2019. Previous studies have shown that these countries have pursued policies compatible with the government IBC. We tested the hypothesis of sustainability of the fiscal stance by examining the cointegration relationship between revenues and expenditures, both as percentages of GDP. The econometric intuition is that if revenues and expenditure can be expressed as a linear combination and residuals can be proven to be stationary, then debt to GDP ratio is mean-reverting, since the difference between revenue and expenditures do not drift wide apart. Hence inferences about long term relationship between revenues and expenditures could be made.

We adopted recent advancements in econometrics to test the fiscal sustainability hypothesis. As a first step, we considered total revenues and total expenditure. Preliminary results indicated that these fiscal variables are not cointegrated and cast doubt on the sustainability hypothesis for the 10 CEEC. The result is also in sharp contrast to earlier panel studies conducted for CEEC, which have all pointed in the direction of cointegrated revenue and expenditures. However, none of the studies considered accounted for structural breaks and cross-sectional dependence in the data generating process, something that has become associated with dynamic macro panels. The study therefore tested, found evidence, and accounted for structural breaks for CEEC – most of which occurred as a result of the changes in fiscal policies prior to joining the EU and also shocks due to business cycles, notably the global financial crises in 2008.

As a next step, the study makes a justification for using cyclically adjusted revenues and expenditures and argues that this represents the long-term discretionary action of the fiscal authorities. Hence, the action of fiscal authorities should be judged by variables which are devoid of business cycle fluctuations or shocks. This is plausible because shocks to fiscal variables induce an automatic response by policymakers and do not necessarily characterise discretionary policy. We use the recently formulated Hamilton filter, which addresses the limitations of the popular HP filter to obtain cyclically adjusted fiscal variables. Results indicate that cyclically adjusted revenue and expenditures are cointegrated with a slope less than unity. We employed the Wald test to ascertain if, indeed, the slope coefficient is unity by way of hypothesis testing since the values are close enough to unity. Results provide enough evidence to reject the null hypothesis of a unit slope coefficient, indicating that the coefficient lies between 0 and 1. Considering the fact that these variables are ratios to GDP, a unit slope of the cointegration is necessary to guarantee strong sustainability in the sense of Quintos (1995). But even though there is cointegration between cyclically adjusted revenue and expenditure, a slope coefficient less than unity implies that expenditures to GDP ratio will grow faster than revenues to GDP ratio, implying a weaker form of sustainability. This is because the debt to GDP ratio is not bounded and therefore not finite. If this continues to happen for a long time, it will generate spikes in the debt to GDP ratio and the fiscal stance will no longer be sustainable.

The possible policy implications are as follows. Firstly, holders of government bonds could lose confidence if debt accumulation is persistent, since this casts doubt with regards to the ability of the government to service its payment. Secondly, the government may have difficulties in marketing its debts to new investors and hence would not be able to raise substantial additional revenue by issuing bonds in the future due to unattractiveness of its debts. Otherwise, government would have to pay high interest in order to make its debt attractive to investors. CEEC governments may therefore have to alter their fiscal policy by way of increasing revenue or reducing expenditure or both as a way of counteracting the deficit problem. The study provides fresh evidence using cyclically adjusted revenue and expenditure for panel sustainability analysis in the context of CEEC. The discretionary action of the government is deemed not to be sufficient to infer strong sustainability of the fiscal stance. The government in CEEC must therefore do more to address the fiscal deficit problem by way of fiscal consolidation to avoid future implications of sustainability.

With the current corona pandemic, fiscal sustainability has become even more challenging as the current recession necessitates further action of the government in terms of stimulating aggregate demand. However, with low revenues due to low productivity and output, government cannot respond adequately to the pandemic without, for instance, borrowing to augment its revenue. Others have also advocated for taxing the super-rich in society as a way of increasing revenue. However, the effectiveness of this policy, as demonstrated by Scheuer and Slemrod (2019), depends on the elasticity of the taxpayers. The current recession and the previous (global financial crises) have taught us that the possibility of a looming recession in the future cannot be ruled out; hence, there should be adequate fiscal space for governments to respond appropriately to future shocks. It is therefore important for government with high debt burdens to institute structural changes, especially in normal times, as a way of reducing debt stocks. This will ensure that there is enough fiscal space in the future to combat the consequences of recessions.

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Appendix

Table A1. Panel stationarity test – first difference of expenditure

Countries	Constant			Constant and trend		
	$k = 1$	$k = 2$	$k = 3$	$k = 1$	$k = 2$	$k = 3$
Czechia	0.195**	0.170	0.132	0.059**	0.094	0.093
Estonia	0.032	0.033	0.065	0.031	0.035	0.031
Hungary.	0.078	0.091	0.110	0.048*	0.089	0.107
Latvia	0.057	0.126	0.070	0.046	0.050	0.046
Lithuania	0.078	0.148	0.080	0.053*	0.097	0.060
Poland	0.075	0.044	0.055	0.045	0.038	0.051
Slovakia	0.122	0.136	0.137	0.047	0.083	0.088
Slovenia	0.064	0.131	0.097	0.060**	0.111*	0.096
Bulgaria	0.069	0.114	0.123	0.031	0.091	0.096
Romania	0.053	0.069	0.181	0.052*	0.040	0.068
Panel statistic	0.970	−0.827	−1.336	4.437***	1.674**	1.034
	(0.166)	(0.796)	(0.909)	(0.000)	(0.047)	(0.151)

Fourier panel stationarity test for 10 CEEC under the Null hypothesis of stationarity.

Critical values (obtained from Becker et al., 2006, p. 289) for individual test statistics are as follows: 0.1318 (10%), 0.1720 (5%), 0.2699 (1%) for $k = 1$; 0.3150 (10%), 0.4152 (5%), 0.6671 (1%) for $k = 2$; 0.3393 (10%), 0.4480 (5%), 0.7182 (1%) for $k = 3$.

Critical values for constant and trend are as follows: 0.0471 (10%), 0.0546 (5%), 0.0716 (1%) for $k = 1$; 0.1034 (10%), 0.1321 (5%), 0.2022 (1%) for $k = 2$; 0.1141 (10%), 0.1423 (5%), 0.2103 (1%) for $k = 3$. Significance at 10%, 5% and 1% are denoted by *, ** and ***, respectively.

CEEC, Central and Eastern European Countries.

Table 14A2. Panel stationarity Test test -– First first difference of Revenuerevenue

Countrs	Constant			Constant and trend		
	$k = 1$	$k = 2$	$k = 3$	$k = 1$	$k = 2$	$k = 3$
Czechia	0.127	0.075	0.074	0.044	0.057	0.074
Estonia	0.093	0.082	0.117	0.051*	0.071	0.111
Hungary.	0.038	0.151	0.159	0.031	0.148**	0.153**
Latvia	0.063	0.062	0.107	0.047	0.052	0.056
Lithuania	0.081	0.122	0.079	0.061**	0.082	0.077
Poland	0.088	0.128	0.157	0.057**	0.056	0.060
Slovakia	0.108	0.531**	0.423*	0.076***	0.120*	0.099
Slovenia	0.061	0.108	0.117	0.057**	0.101	0.077
Bulgaria	0.180**	0.327	0.366*	0.027	0.149**	0.159**
Romania	0.038	0.065	0.083	0.035	0.048	0.041
Panel statistic	1.282	0.574	0.292	4.723***	2.965***	2.357***
	(0.100)	(0.283)	(0.385)	(0.000)	(0.002)	(0.009)

Fourier panel stationarity test for 10 CEEC under the Null hypothesis of stationarity.

Critical values (obtained from Becker et al., 2006, p. 289) for individual test statistics are as follows: 0.1318 (10%), 0.1720 (5%), 0.2699 (1%) for $k = 1$; 0.3150 (10%), 0.4152 (5%), 0.6671 (1%) for $k = 2$; 0.3393 (10%), 0.4480 (5%), 0.7182 (1%) for $k = 3$.

Critical values for constant and trend are as follows: 0.0471 (10%), 0.0546 (5%), 0.0716 (1%) for $k = 1$; 0.1034 (10%), 0.1321 (5%), 0.2022 (1%) for $k = 2$; 0.1141 (10%), 0.1423 (5%), 0.2103 (1%) for $k = 3$. Significance at 10%, 5% and 1% are denoted by *, ** and ***, respectively.

CEEC, Central and Eastern European Countries.

Table 15A3. Panel Stationarity stationarity test with sharp breaks - – Robustness robustness check

Panel - – Panel test (revenue)		
Model	**Level**	**First difference**
Break (Homogeneous)	1.762 (0.039)**	0.426 (0.335)
Breaks (Heterogeneous)	2.975 (0.001)***	0.152 (0.440)
Panel B – Panel stationarity test (expenditure)		
Model	**Level**	**First difference**
Break (Homogeneous)	1.687 (0.046)**	0.137 (0.446)
Breaks (Heterogeneous)	2.031 (0.021)**	0.630 (0.264)

Panel test by Carrion-i-Silvestre et al. (2005). Reported test statistics and *P*-values in parenthesis. *, ** and *** indicate rejection of the null hypothesis of 'stationarity' at 10%, 5% and 1%, respectively.

Table A4. Panel stationarity test – cyclically adjusted expenditure

Countries	**Constant**			**Constant and trend**		
	$k = 1$	$k = 2$	$k = 3$	$k = 1$	$k = 2$	$k = 3$
Czechia	0.144*	0.165	0.132	0.065**	0.134**	0.126**
Estonia	0.080	0.401	0.385*	0.046	0.105*	0.092
Hungary.	0.157*	0.096	0.068	0.059**	0.100	0.126**
Latvia	0.106	0.202	0.155	0.045	0.080	0.073
Lithuania	0.089	0.172	0.146	0.072***	0.135**	0.126**
Poland	0.154*	0.466**	0.577**	0.047	0.074	0.073
Slovakia	0.095	0.258	0.206	0.063**	0.146**	0.126**
Slovenia	0.091	0.488**	0.332	0.070**	0.140**	0.093
Bulgaria	0.102	0.072	0.097	0.068**	0.072	0.065
Romania	0.064	0.087	0.110	0.049*	0.089	0.110
Panel statistic	2.491***	2.376***	1.461*	7.017***	4.511***	2.885***
	(0.006)	(0.009)	(0.072)	(0.000)	(0.000)	(0.002)

Fourier panel stationarity test for 10 CEEC under the Null hypothesis of stationarity.

Critical values (obtained from Becker et al., 2006, p. 289) for individual test statistics are as follows: 0.1318 (10%), 0.1720 (5%), 0.2699 (1%) for $k = 1$; 0.3150 (10%), 0.4152 (5%), 0.6671 (1%) for $k = 2$; 0.3393 (10%), 0.4480 (5%), 0.7182 (1%) for $k = 3$.

Critical values for constant and trend are as follows 0.0471 (10%), 0.0546 (5%), 0.0716 (1%) for $k = 1$; 0.1034 (10%), 0.1321 (5%), 0.2022 (1%) for $k = 2$; 0.1141 (10%), 0.1423 (5%), 0.2103 (1%) for $k = 3$ Significance at 10%, 5% and 1% are denoted by *, ** and ***, respectively.

CEEC, Central and Eastern European Countries.

Table 18A5.: Panel stationarity Test test -– Cyclically cyclically adjusted revenue

Countrs	Constant			Constant and trend		
	$k=1$	$k=2$	$k=3$	$k=1$	$k=2$	$k=3$
Czechia	0.160*	0.416**	0.519**	0.048*	0.070	0.060
Estonia	0.079	0.414*	0.249	0.044	0.042	0.089
Hungary	0.040	0.398*	0.348*	0.039	0.058	0.103
Latvia	0.037	0.324*	0.356*	0.037	0.036	0.051
Lithuania	0.061	0.242	0.165	0.059**	0.145**	0.111
Poland	0.241**	0.434**	0.435*	0.078***	0.056	0.072
Slovakia	0.080	0.262	0.230	0.056**	0.149**	0.123*
Slovenia	0.097	0.237	0.261	0.075***	0.132*	0.066
Bulgaria	0.040	0.281	0.343*	0.037	0.035	0.071
Romania	0.083	0.246	0.338	0.054*	0.044	0.073
Panel statistic	1.524***	4.395***	3.768***	5.575***	1.985**	1.682**
	(0.064)	(0.000)	(0.000)	(0.000)	(0.023)	(0.046)

Fourier panel stationarity test for 10 CEEC under the Null hypothesis of stationarity.

Critical values (obtained from Becker et al., 2006, p. 289) for individual test statistics are as follows: 0.1318 (10%), 0.1720 (5%), 0.2699 (1%) for $k = 1$; 0.3150 (10%), 0.4152 (5%), 0.6671 (1%) for $k = 2$; 0.3393 (10%), 0.4480 (5%), 0.7182 (1%) for $k = 3$.

Critical values for constant and trend are as follows 0.0471 (10%), 0.0546 (5%), 0.0716 (1%) for $k = 1$; 0.1034 (10%), 0.1321 (5%), 0.2022 (1%) for $k = 2$; 0.1141 (10%), 0.1423 (5%), 0.2103 (1%) for $k = 3$.

Significance at 10%, 5% and 1% are denoted by *, ** and ***, respectively.

CEEC, Central and Eastern European Countries.

Table A6. Panel stationarity test – first difference of cyclically adjusted expenditure

Countries	Constant			Constant and trend		
	$k=1$	$k=2$	$k=3$	$k=1$	$k=2$	$k=3$
Czechia	0.224**	0.255	0.298	0.140***	0.142**	0.137*
Estonia	0.041	0.105	0.061	0.037	0.043	0.045
Hungary.	0.059	0.141	0.144	0.045	0.141**	0.131*
Latvia	0.076	0.075	0.068	0.056**	0.063	0.068
Lithuania	0.121	0.211	0.194	0.119***	0.126*	0.118*
Poland	0.193*	0.191	0.198	0.051*	0.120*	0.144**
Slovakia	0.044	0.304	0.204	0.036	0.036	0.063
Slovenia	0.138*	0.189	0.147	0.108***	0.137**	0.074
Bulgaria	0.102	0.073	0.142	0.049*	0.051	0.061
Romania	0.042	0.147	0.055	0.038	0.048	0.044
Panel statistic	2.237**	0.668	−0.086	9.258***	3.130***	2.196**
	(0.013)	(0.252)	(0.532)	(0.000)	(0.001)	(0.014)

Fourier panel stationarity test for 10 CEEC under the Null hypothesis of stationarity.

Critical values (obtained from Becker et al., 2006, p. 289) for individual test statistics are as follows: 0.1318 (10%), 0.1720 (5%), 0.2699 (1%) for $k = 1$; 0.3150 (10%), 0.4152 (5%), 0.6671 (1%) for $k = 2$; 0.3393 (10%), 0.4480 (5%), 0.7182 (1%) for $k = 3$.

Critical values for constant and trend are as follows 0.0471 (10%), 0.0546 (5%), 0.0716 (1%) for $k = 1$; 0.1034 (10%), 0.1321 (5%), 0.2022 (1%) for $k = 2$; 0.1141 (10%), 0.1423 (5%), 0.2103 (1%) for $k = 3$. Significance at 10%, 5% and 1% are denoted by *, ** and ***, respectively.

CEEC, Central and Eastern European Countries.

Table 20A7.: Panel stationarity Test test -– First first difference of cyclically adjusted revenue

Countrs	Constant			Constant and trend		
	$k = 1$	$k = 2$	$k = 3$	$k = 1$	$k = 2$	$k = 3$
Czechia	0.060	0.092	0.168	0.058**	0.068	0.067
Estonia	0.043	0.073	0.097	0.039	0.067	0.086
Hungary.	0.042	0.071	0.121	0.040	0.071	0.122*
Latvia	0.093	0.096	0.092	0.093***	0.096	0.088
Lithuania	0.166*	0.496**	0.303	0.103***	0.130*	0.104
Poland	0.217**	0.187	0.286	0.033	0.071	0.089
Slovakia	0.102	0.244	0.214	0.102***	0.104*	0.095
Slovenia	0.131	0.172	0.185	0.114***	0.197**	0.088
Bulgaria	0.1068	0.085	0.117	0.063**	0.077	0.116*
Romania	0.120	0.071	0.052	0.039	0.047	0.042
Panel statistic	2.244**	0.421	0.190	9.430***	3.306***	2.275**
	(0.012)	(0.337)	(0.425)	(0.000)	(0.000)	(0.011)

Fourier panel stationarity test for 10 CEEC under the Null hypothesis of stationarity.

Critical values (obtained from Becker et al., 2006, p. 289) for individual test statistics are as follows: 0.1318 (10%), 0.1720 (5%), 0.2699 (1%) for $k = 1$; 0.3150 (10%), 0.4152 (5%), 0.6671 (1%) for $k = 2$; 0.3393 (10%), 0.4480 (5%), 0.7182 (1%) for $k = 3$.

Critical values for constant and trend are as follows: 0.0471 (10%), 0.0546 (5%), 0.0716 (1%) for $k = 1$; 0.1034 (10%), 0.1321 (5%), 0.2022 (1%) for $k = 2$; 0.1141 (10%), 0.1423 (5%), 0.2103 (1%) for $k = 3$ Significance at 10%, 5% and 1% are denoted by *, ** and ***, respectively.

CEEC, Central and Eastern European Countries.

Table A8. Panel cointegration test of cyclically adjusted revenue and cyclically adjusted spending (HP Filter used for detrending series)

	$Z_{\tau}(N)$		$Z_{\varphi}(N)$	
Models	**Value (τ)**	***P*-value**	**Value (φ)**	***P*-value**
No breaks	1.731	0.958	−2.409***	0.008
Level break	−1.984**	0.024	−2.490***	0.006
Regime shift	−6.050***	0.000	−7.182***	0.000
Number of observations	250		250	

Westerlund and Edgerton (2008) cointegration test. Maximum of three breaks are permitted. Displayed *P*-values are based on one-sided normal distribution test. *, ** and *** denote rejection of the null hypothesis at 10%, 5% and 1%, respectively. Maximum of three structural breaks in the cointegration relationship. Detrending of the series was done using the HP filter.

HP, Hodrick Prescott.