Essays on Public Debt Sustainability Analysis with Applications from Parametric and Non-Parametric Panel Data models

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Benjamin Kwaw Owusu

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First advisor: Prof. Dr. Alfred Greiner
Second advisor: Prof. Dr. Dietmar Bauer

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Chapter 1

Introduction

"A government that badly mismanages its finances and ends up saddling the state and future generations with an unsustainable debt will not long retain its legitimacy. But neither will a government that stubbornly refuses to borrow to meet an emergency or invest in the future when productive investment opportunities present themselves."

Eichengreen et. al., 2021 p.2,
In Defense of Public Debt

1.1 Overview

This thesis examines the concept of debt sustainability using panel econometric analysis and consists of three main empirical papers. A common feature of debt sustainability analysis connects all three papers. Public debt sustainability analysis entails examining if, indeed, governments have the ability to meet all their current and future financial obligations without default. It requires that the debt-to-GDP ratio remains finite, and moreover, the market should be willing to hold government debt (Wickens, 2008). Debt sustainability analysis is therefore significant in assessing fiscal risk given current and future projections of macroeconomic variables. With these papers, I answer the same research question of whether the public finance (fiscal stance) for a group of countries is sustainable using three different econometric approaches involving both parametric and non-parametric econometric models.

In the first paper, debt sustainability analysis is conducted for Central and Eastern European countries through the lens of cointegration analysis of revenues and expenditures. By investigating the long-run relationship between revenues and expenditures using panel cointegration analysis, we are able to make inferences about the sustainability of the fiscal stance according to Hakkio and Rush (1991), Quintos (1995) and Afonso (2005). The paper employs a panel econometric technique that accounts for structural breaks and cross-sectional dependence in the cointegration regression. Following Gali et al. (2003) and Blanchard (2006), we use the structural component of revenues and expenditures for the cointegration analysis since they represent the long-term behaviour of fiscal authorities or policy makers.

The second paper analyzes debt sustainability for a group of EU countries by estimating a fiscal reaction function. We estimated the response of the primary balance-to-GDP ratio (fiscal policy variable) to changes in the debt-to-GDP ratio and other macroeconomic control variables which influence fiscal policy. In addition to the traditional linear reaction function proposed by Bohn (1995,1998), we also study the nonlinear behaviour of the primary balance to variations in the debt-to-GDP ratio using panel splines (a non-parametric or semi-parametric estimator). The approach enables us to visualize the response of fiscal policy variable to the distribution of the debt-to-GDP ratio. We study two main clusters of countries based on categorization by a data-driven algorithm (k-means clustering). We obtained two main clusters: a cluster comprising big economies (with high debt comparatively) and a cluster dominated by smaller economies. We also study sub-samples consisting of financial or economic crisis periods and the non-crisis periods.

Regarding the third paper, we used a similar dataset and estimated a fiscal reaction function considering nonlinearities in the reaction coefficients (slope parameters) to obtain multiple regimes based on the heterogeneity in the coefficient of the transition function. Specifically, we resort to a regime-switching model where changes between regimes occur smoothly via a logistic regression function. Hence, we are able to estimate fiscal policy behaviour separately for each regime based on a certain transition variable. Thus, we studied fiscal policy actions in a regime of low debt and a regime of high debt. Additionally, we considered the concept of fiscal space and its effect on fiscal policy by estimating a fiscal reaction function for regimes of low and high fiscal space.

1.2 Chapters and main findings

The next subsections present a summary of each of the three essays. The motivation, methodology, data and findings for each paper are briefly discussed.

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

Before the accession to the European Union (EU), governments in most Central and Eastern European 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 is to ensure that they meet the necessary fiscal criteria regarding size of debt, deficit and other obligations stipulated in the Maastricht Treaty (MT) and Stability and Growth Pact (SGP). Eight CEECs out of the ten countries that joined the EU from the Eastern enlargement scheme had debt-to-GDP ratios below the 60% threshold required by the MT and SGP. Hence, Hallett and Lewis (2007) speculated that these CEECs could follow a high debt path for years without violating the fiscal sustainability requirements. Several years after joining the EU, whether these countries have pursued sustainable fiscal policies remains to be seen. Previous panel sustainability studies focusing on CEECs did not account for the possibility of structural breaks and cross-sectional dependence in the cointegration relationship. There is empirical evidence that ignoring structural changes and cross-sectional dependence could affect the credibility of econometric inferences (for instance, see Westerlund and Edgerton (2008). Bai and Perron (1998) and Carrion-i-Silvestre et al. (2005)). The accession of member states to the EU could, for instance, represent a structural change due to requirements that must be met and maintained by union members, notably requirements enshrined in the MT and SGP. Furthermore, the fact that countries in the sample are linked geographically and economically implies that these countries could respond to shocks similarly. Also, the possibility of spillover effects between countries is high; therefore, the cross-sectional dependence issue in the panel data could be prevalent. Hence, accounting for structural breaks and cross-sectional dependence in the cointegration process is significant for inference and, therefore, policy conclusion.

This paper assesses the fiscal sustainability hypothesis for 10 CEECs between 1997 and 2019. The study uses recent panel data econometric methods where we account for structural breaks and cross-sectional dependence in the cointegration relationship between government revenues and expenditures. Initial results did not find evidence of cointegration, and hence no support for sustainability of the fiscal stance. Further, we distinguish between structural and cyclical components of revenues and expenditures to emphasize the structural component. Our conjecture is that the structural component of fiscal variables represents the actual long-term behaviour of the fiscal authorities. Using structural revenues and expenditures (cyclically adjusted), the study found lack of evidence of cointegration between structural revenues and expenditures. Expenditures and revenues are deemed not to have a long-run stable relationship, and if this continues for a long time, the government may find it difficult to market its debt in the long-run. The result suggests that the fiscal authorities in CEECs must therefore, take long-term actions to counteract the rising fiscal deficit problems.

A similar variant of this essay has been published as Owusu, B. (2021), "Fiscal Sustainability Hypothesis Test in Central and Eastern Europe: A Panel Data Perspective", *Central European Economic Journal*, vol. 8 (55), pp. 285-312.

The following are the changes that distinguish this thesis from the published version. Firstly, a summary description of the cointegration test has been provided in the main text, which is not present in the journal's published version. This is gives the reader some background and understanding of the cointegration test. Secondly, I replaced the Fourier unit root test (from Nazlioglu and Karul (2017)) by a standard panel data unit root test, which entails both the first generational unit root test (Im Pesaran and Shin test (IPS) and Levin, Li and Chu test (LLC)) and second generational unit root test (Cross-sectionally augmented Im Pesaran and Shi test (CIPS)) that account for cross-sectional dependence. The implementation and interpretation of the proposed unit root test are non-complex and relatively straightforward. Lastly, regarding the cointegration test, only two (and one) structural breaks are considered in this thesis which is justifiable because the number of time series observation is not long enough.

Assessing nonlinearities and heterogeneity in debt sustainability analysis: A panel spline approach

This paper empirically studies public debt sustainability analysis using panel penalized splines for 25 EU economies from 2000 to 2019 to estimate a fiscal reaction function of the primary balance to changes in the lagged debt-to-GDP ratio.

The standard assumption in debt sustainability analysis (see Bohn (1995,1998)) assumes a linear relationship between the response variable and the covariates and allows us to measure the marginal effects of the latter on the former. However, in this paper, we intend to get additional insights by going beyond the standard linear fixed effects model. The hypothesis that the fiscal reaction coefficient in terms of the response of the discretionary fiscal policy is constant or uniform across the distribution of debt is questioned. Instead, we allow it to change with the magnitude of the debt ratios. We show that the reaction of the structural primary balance changes with the size of the debt ratio, implying that, depending on the size of the debt (low or high debt ratios), the relationship between the primary balance and the debt ratio is expected to be different. Using data from the AMECO European Commission, we consider two clusters of EU countries and distinguish between good or normal economic times and crisis times. The reason for focusing on this sample is against the background of the sovereign debt crisis in Europe around the year 2010.

A positive coefficient, on average, indicates sustainable policies, which is supported by the most of our estimates. Moreover, we show that this relationship is not homogeneous across the distribution of the debt ratios but varies with the magnitude of public debt-to-GDP. The estimations for a cluster of mainly big economies reveal a strongly increasing (steep) reaction for small and high debt ratios, while it is rather less steep for intermediate levels. For the sub-sample of economic crisis period, the reaction coefficient increases initially, flattens at relatively high debt and then trends upwards at very high debt levels. Additionally, for a cluster consisting of smaller EU economies and for a period of normal economic times, there is an indication of 'fiscal fatigue', meaning that the reaction of the primary balance becomes weaker and eventually falls as the debt-to-GDP ratio exceeds some threshold.

Thus, our estimations provide support for the hypothesis of heterogeneity in the data. The results are significant for policy implications as they indicate that the size of the current debt and the economic situation are essential for the assessment and evaluation of fiscal sustainability. Our results show that size does matter indeed. The level of the explanatory variable influences the fiscal reaction coefficient. Therefore, policy recommendations need to consider the status of the current debt situation to be successful, as the reactions show different behaviour for low debt levels compared to medium or high ratios.

A similar variant of this essay has been published as Owusu, B., Bökemeier, B. Greiner, A. (2023), "Assessing nonlinearities and heterogeneity in debt sustainability analysis: a panel spline approach," *Empirical Economics*, 64, pp. 1315–1346.

The following are changes implemented that differentiate the thesis's content from the published version in the journal. Firstly, standard econometric tests are implemented before model specification. For instance, I did not assume that the slope coefficient of the econometric model could be pooled but rather implemented a test to that effect. Hence, the specified model is augmented with dummy variables to enable pooling. Secondly, I tested for the presence of individual and time effects in the data. Finally, the estimation of standard errors in this paper accounted for serial correlation, heteroskedasticity and cross-sectional dependence problems. For the non-parametric model, residuals are inspected to detect cross-sectional dependence.

Regime-based debt sustainability analysis: Evidence from Euro area economies

This paper empirically studies nonlinearities in debt sustainability analysis using the estimation technique of panel smooth transition regression (PSTR). We assess Euro area debt sustainability by analyzing a fiscal reaction function of changes in the cyclically adjusted primary balance in response to variations in the debt-to-GDP ratio. Applying PSTR for panel data in annual frequency from 2000-2019, we estimate the existence of a threshold in the behaviour of the reaction function, refrain from the country-wise perspective and apply a regime-switching model to model nonlinearities. Data is segregated into different regimes endogenously via a logistic regression function.

The distinctions between low and high debt regimes also reveal heterogeneous behaviour of the cyclically adjusted primary balance across the distribution of the debt-to-GDP ratios. The coefficients are positive for both regimes; however, the coefficient in the low-debt regime is not statistically significant. For a sub-sample of highly indebted countries, we find a statistically significant negative (positive) reaction coefficient for the low (high) debt regime. Thus, debt sustainability seems to be given in the high debt regime. Several robustness tests support our findings.

Further, we explored the role of fiscal space in the context of debt sustainability analysis. We first determined the fiscal space for each euro area economy, with fiscal space is defined as the difference between the reaction coefficient and the interest-growth differential. The average value of the fiscal space turned out to be strictly positive, implying that the debt-to-GDP ratio remains bounded, except for Greece and Cyprus. In the next step, we estimated the reaction coefficients for different regimes of the fiscal space. When we distinguish between 3 groups of countries, high, medium and low public debt, we obtained a mixed picture for the low fiscal space regime. In contrast, in the regime with high fiscal space the reaction turned out to be positive and statistically significant, independent of the estimated model. When no distinction is made, both the estimations for the low fiscal regime and those for the high fiscal regime yield a positive and statistically significant reaction of the primary surplus to higher debt, relative to GDP, respectively. Finally, we saw that in situations of high fiscal space, the effect of crisis is less pronounced than in situations of low fiscal space. This is probably because economies with higher fiscal space have the capacity to implement appropriate fiscal interventions to dampen the effects of crisis more easily compared to economies with a low fiscal space without jeopardizing fiscal sustainability. Our study contributes to the literature of the application of panel data for debt sustainability studies using a panel smooth transition regression model

A similar variant of this essay has been published as Owusu, B., Bökemeier, B., Greiner, A. (2023), "Regime-based debt sustainability analysis: Evidence from euro area economies", *European Journal of Political Economy,* 102458, ISSN 0176-2680, https://doi.org/10.1016/j.ejpoleco.2023.102458.

The changes implemented in the thesis are as follows. I have included a subsection where I discussed the concept of fiscal space and examined its role in debt sustainability analysis. Fiscal space is used as a transition variable in the PSTR framework. I also estimated a robustness test considering a model with both individual and time effects. Finally, I visualized the residual plot to investigate potential cross-sectional dependence in the residuals.

1.3 Contributions

Two of the three chapters in this thesis are co-authored. Chapter one is my single authorship paper. Chapters two and three are joint works with Prof. Greiner and Dr. Bökemeier. In both chapters, I contributed to the conceptual idea, the empirical aspect of the paper (I performed all the econometric estimations) and the analysis of results.

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Chapter 2

Fiscal sustainability hypothesis test in central and eastern Europe: A panel data perspective

2.1 Introduction

The recent financial crises and global economic downturn prompted governments' interventions worldwide through fiscal expansions in attempts to stimulate aggregate demand. This has implications for fiscal policy since government has to find means of financing its deficits. Rising public deficits and debts to unsustainable levels may have long-run implications for the government since holders of government debts could lose confidence in holding government bonds. Secondly, the government could also default on its debts if it reaches unsustainable levels. The need to finance the public deficit imposes a constraint on fiscal policy since governments in dynamically efficient economies have borrowing limits and face a present value borrowing constraint.[1] The issue of fiscal sustainability has, therefore, received considerable attention both in theoretical and empirical discussions.

Prior to the accession of the European Union (EU), governments in Central and Eastern European Countries (CEECs) 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 is to ensure that they meet the necessary fiscal criterion in terms of the size of debts, deficits and other obligations as stated in the Maastricht Treaty (MT) and Stability and Growth Pact (SGP). Eight CEECs out of the ten countries that joined the EU from the so-called Eastern enlargement scheme had lower debt-to-GDP ratios below the 60% threshold required by the MT and SGP. Hence, Hallett

[1] See Abel et al. (1988) for a detailed discussion regarding the dynamic efficiency of an economy.

and Lewis (2007) speculated that these CEECs could follow a high debt path for years without necessarily violating the fiscal sustainability requirements. Several years after joining the EU, it remains to be seen if these countries have indeed 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), who found that US debt and deficit are stationary using annual data from 1962 to 1984. Trehan and Walsh (1988) tested and confirmed the stationarity of US public deficit and debt (for data between 1960 and 1984). Hakkio and Rush (1991) investigated the cointegration of revenues and expenditure inclusive of interest payment. They concluded that the US government budget deficit is too large and not consistent with fiscal sustainability. Tanner and Lui (1994) also investigated the cointegration between US government revenue and expenditure (between 1950 and 1989). They revealed that there is evidence of cointegration and, hence, support for fiscal sustainability if structural break is accounted for. Similarly, Quintos (1995) examined the cointegration between US government revenue and expenditure and concluded that the government fiscal deficit had followed a sustainable path. Using US (1692-1992) and UK (1792-1992) data, Ahmed and Rogers (1995) tested the cointegration between revenues and expenditure and confirmed the sustainability of the fiscal stance for both countries. Afonso (2005) performed a cointegration analysis for 15 individual EU countries for data between 1970-2003 and found that with a few exceptions, most EU countries had followed an unsustainable fiscal policy path. All these papers laid the foundation for the cointegration approach between government revenues and expenditures to test the fiscal sustainability analysis.

Even though a vast stream of empirical studies on fiscal sustainability in the European continent have been undertaken, there exist limited studies in the context of CEECs (Bökemeier and Stoian, 2018). Among the fewer studies includes Krajewski et al. (2016), who examined the public debt sustainability for ten selected CEECs using panel stationarity, a cointegration technique and a fiscal reaction function from 1990-2012. Their results indicated that the fiscal stance of selected CEECs is jointly sustainable. Similarly, Llorca and Redzepagic (2008) assessed the sustainability of fiscal policy for eight CEECs using panel cointegration analysis and found that these countries pursued sustainable fiscal policies for the period from 1999 to 2006 using quarterly data. Bökemeier and Stoian (2018) also investigated debt sustainability in 10 CEECs using estimates of a fiscal reaction function in its cubic form from 1998 to 2015. Their results revealed that government debts were at sustainable levels and that governments had not reached fiscal fatigue thresholds.[2]. Even though the studies above employed panel sustainability test for CEECs, they did not incorporate the possibility of structural breaks and cross sectional dependence in the panel data cointegration relationship[3].

The paper aims to ascertain the fiscal sustainability of 10 CEECs for the period 1997 to 2019 by investigating the long-run relationship between revenues and expenditure using panel cointegration.[4] The study employs a panel data analysis 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. The possibility of breaks in the data is therefore high (Carrion-i-Silvestre et al., 2005). In cointegration analysis, structural changes have the tendency of affecting the cointegration vector and, hence, affects the long-run stable relationship between the variables (Westerlund and Edgerton, 2008). Bai and Perron (1998) and Carrion-i-Silvestre et al. (2005) both show that ignoring structural changes (breaks) in the data-generating process could lead to wrong inferences in econometric analysis. Furthermore, cross-country macroeconomic and financial datasets are associated with cross-sectional dependence because of inter-country links and dependencies (Westerlund and Edgerton, 2008). Cross-sectional dependence affects the size properties of the unit root test, for instance, and hence, affects the credibility of inferences.[5] This study, therefore, adopts the so-called "second generational" econometric procedure which accounts for cross-sectional dependence in the unit root test and cointegration relationship.

The preliminary result of the study shows that revenues and expenditures do not have a long-term relationship and, hence, a rejection of the sustainability hypothesis. Further, we discriminate between structural and cyclical components of revenues and expenditures to emphasise the structural component. This is another novelty of this paper when compared to 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), 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 the business cycle represents discretionary fiscal policy, we consider structural fiscal variables which provide a benchmark by which fiscal policy can be judged (Blanchard, 2006). We argue that these structural components of fiscal variables (cyclically adjusted

[2] Fiscal fatigue is a terminology introduced by Gosh et al. (2013), which refers to the situation where debt-to-GDP reaches a threshold such that it becomes increasingly difficult to generate a higher primary balance in response to higher debt so that an inverse relationship ensues due to increased burden on fiscal policy

[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 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 could be lower than the empirical size.

variables) represent the actual long term behaviour of policymakers and should be considered when conducting sustainability analysis.

Further results (when cyclically adjusted variables are considered) point to the lack of strong evidence of cointegration between revenue and expenditure. Hence, the debt-to-GDP ratio is not bounded, and the debt stock is considered not to be finite or sustainable. 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 threefold. Firstly, it employs recent advances in panel econometrics that simultaneously model structural breaks and cross-sectional dependence in the cointegration regression for the sustainability hypothesis test (cointegration analysis) for CEEC data. Secondly, the study makes a case for using structural fiscal variables devoid of automatic response in the cointegration analysis for the sustainability hypothesis test. Finally, the study adds to the growing literature on fiscal sustainability for panel data in the CEECs region.

The rest of the study is structured as follows. Section 2.2 will discuss the methodology by laying the theoretical foundations for the sustainability test. Section 2.3 will further discuss the econometric tests, estimations and discussion of results. Section 2.4 will finally conclude the paper.

2.2 Methodology

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

$$g_t + i_t b_{t-1} = r_t + b_t - b_{t-1} \tag{2.1}$$

g_t represents government spending, b_t is the stock of 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 on the Left Hand Side (LHS) inclusive of interest payment is equal to total government receipts at the Right Hand Side (RHS) at all times in order for the budget constraint to hold intertemporarily. Here, we rule out the possibility of monetization of debt. 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 and Fincke, 2015) with little or no influence from fiscal authorities.

Taking the state of the economy into consideration, (2.1) can be re-written as

$$\frac{g_t}{y_t} + \frac{(1+i_t)b_{t-1}}{y_t} = \frac{r_t}{y_t} + \frac{b_t}{y_t} \tag{2.2}$$

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

$$\frac{g_t}{y_t} + \frac{(1+i_t)b_{t-1}}{y_t} = \frac{r_t}{y_t} + \frac{b_t}{y_t} \tag{2.3}$$

$$\frac{g_t}{y_t} + \frac{(1+i_t)}{(1+f_t)}.\frac{b_{t-1}}{y_{t-1}} = \frac{r_t}{y_t} + \frac{b_t}{y_t} \tag{2.4}$$

where f is the nominal growth rate of the economy (GDP). Using capital notations, (2.4) can be rewritten as

$$G_t + (1+\rho_t)B_{t-1} = R_t + B_t \tag{2.5}$$

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

Further modification is required for empirical estimation; let $E_t = G_t + (\rho_t - \rho)B_{t-1}$ be the government expenditure inclusive of interest rates where ρ is the mean of ρ_t assumed to be stationary. Assuming that (2.5) holds continuously at each time t, then by forward substitution the present value budget constraint can be written as[6]

$$B_{t-1} = \sum_{s=t+1}^{\infty} \left(\frac{1}{1+\rho}\right)^{s-t} (R_{s-1} - E_{s-1}) + lim_{s\to\infty} \left(\frac{1}{1+\rho}\right)^{s-t} B_{s-1} \tag{2.6}$$

Sustainability implies that the second term on the RHS of (2.6) converges to zero as time approaches infinity. This is also known as the transversality condition, which constrains the debt-to-GDP 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 current and future discounted primary surpluses. As pointed out by Afonso (2005), the "no Ponzi" condition can be assessed empirically by testing the stock of debt for stationarity. Earlier studies that focused on testing the stationarity of public debt include Kremers (1988), Wilcox (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). Assuming a stationary real interest rate and applying the difference operator, the present value budget constraint can be re-written as

[6] Proof of 2.6 can be found in subsection 2.5.1 in the appendix.

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

$$GG_{t-1} - R_{t-1} = \sum_{s=t+1}^{\infty} \left(\frac{1}{1+\rho}\right)^{s-t} (\Delta R_{s-1} - \Delta E_{s-1}) + lim_{s\to\infty} \left(\frac{1}{1+\rho}\right)^{s-t} \Delta B_{s-1} \tag{2.7}$$

GG_t is given by $G_t + i_t B_{t-1}$. Testing for the sustainability hypothesis in this paper, we assume "no ponzi" scheme, which implies that the second term on the RHS of 2.7 approaches zero as time approaches infinity (following Afonso (2005)). Next, we assume R_t and E_t are non-stationary at their levels but stationary at their first difference. The LHS (GG_t - R_t) must also be stationary for (2.7) to hold. This implies testing for the difference stationarity for GG_t and R_t. It 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, 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 other words, these variables' differences do not drift wide apart. Hence, they are cointegrated because they have a long-run stable relationship. From (2.7), this implies testing if GG_t and R_t are integrated of order 1 ($I(1)$) with cointegration vector (1, -1) as argued by Quintos (1995). One can test for cointegration equivalently as below:

$$R_t = \alpha + \gamma GG_t + \mu_t \tag{2.8}$$

Where μ_t is a zero mean stationary process. Alternatively making use of the expression $\Delta B_t = GG_t - R_t$, then from 2.8 we get

$$\Delta B_t = (1-\gamma)GG_t - \alpha - \mu_t \tag{2.9}$$

Hakkio and Rush (1991) argued that the $0 < \gamma \leq 1$ guarantees sustainability if variables are cointegrated. However, given that variables are expressed as ratios of GDP, $\gamma = 1$ implies sustainability since from (2.9), debt-to-GDP ratio is bounded and will grow at a constant rate (Afonso, 2005). It is important to make a special remark about the condition $0 < \gamma < 1$. At this point the government expenditure exceeds its revenue and, therefore, the possibility of default could be high if this continues for a long time. It will be difficult to market its bonds and the government may have to pay high interest rates in order 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 whilst γ less than 1 implies some weaker form of sustainability. For $\gamma > 1$, revenues are strictly higher than expenditure; hence, sustainability of the fiscal stance is not an issue (Afonso, 2005).

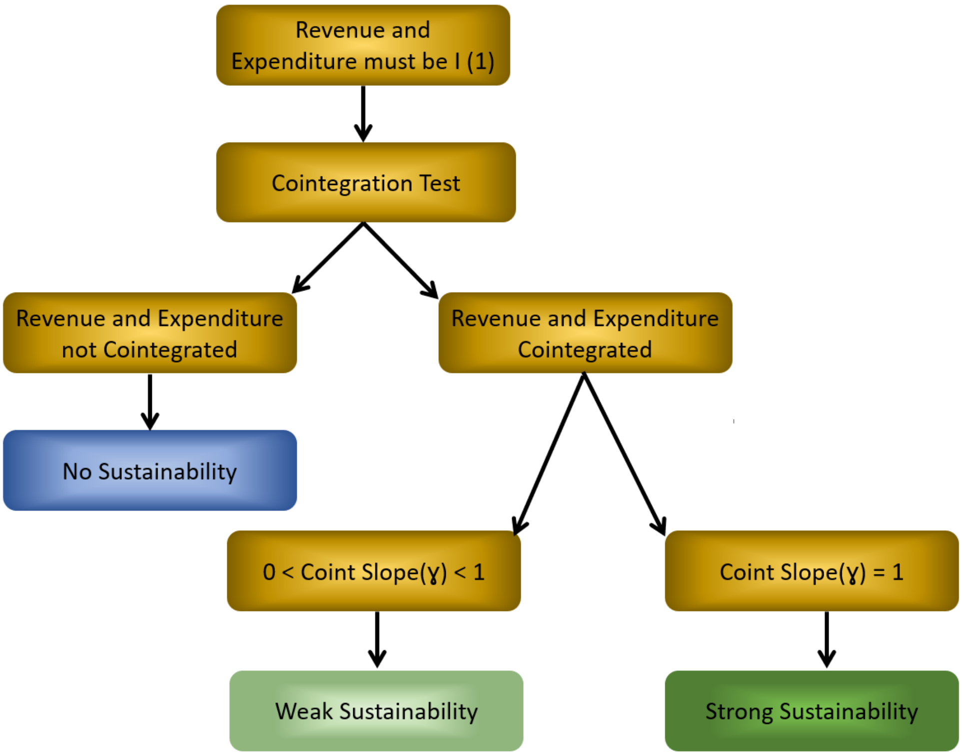
Empirical cointegration test for (2.8) can be conducted conventionally by regressing R_t on GG_t by OLS and testing the residuals for stationarity to confirm if cointegration holds. Westerlund and Prohl (2010) argued that such a conventional test fails to reject the null hypothesis of no cointegration very often, which implies a rejection of the 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, provide more variability, less collinearity among variables and give 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. Finally, panel analysis helps to control for the effects of omitted variables bias in econometrics (Hsiao, 2003). Hence, the study will resort to a panel test which will subject the residuals in (2.8) to a cointegration test. The test is flexible enough to account for structural breaks and cross-sectional dependence, which is common to panel data analysis. Figure 2.1 provides a graphical overview of the cointegration procedure for testing fiscal sustainability. The procedure stipulates that the variables must be integrated of order 1 if they are not stationary at their levels. A conintegration test can then be carried out if this condition is met. As shown by Afonso (2005), if variables are not cointegrated, there is no fiscal sustainability.[8] However, if cointegration exists for the variables, there is the evidence of fiscal sustainability and therefore we proceed to ascertain the strength of the fiscal sustainability by estimating the slope parameter. A slope coefficient less than 1 implies a weak form of sustainability (unbounded debt-to-GDP ratio) whilst a coefficient equal to 1 implies strong sustainability.

2.3 Empirical tests, estimations and results

Firstly, we present a review of past empirical papers on fiscal sustainability with a focus on panel studies. Subsequently, we will discuss the datasets and some data characteristics, after which we shall proceed with the empirical test of the fiscal sustainability. Table 2.1 shows previous papers on panel data fiscal sustainability analysis for mostly CEECs, Organisation for Economic Corporation and Development (OECD), and EU countries. Regarding CEECs, previous studies notably by Llorca and Redzepagic (2008), Krajewski et al. (2016) and Bökemeier and Stoian (2018), all point to the direction of a sustainable fiscal policy. It will therefore be interesting to compare our results directly to their studies.

[8] There is an exceptional case where revenue grows faster than expenditure. In that case, even though the two variables may not necessarily be cointegrated, fiscal sustainability is not a problem (Afonso, 2005).

Figure 2.1: Cointegration procedure for fiscal sustainability test



Source: Constructed by author

Regarding our dataset, revenue and expenditure variables were all obtained from the OECD website (OECD, 2020) for the 10 CEECs.[9] All data exists in annual frequency. The sample period is from 1995 to 2019. A total of 250 observations is generated from a combination of 10 countries over a 25 year period. It is important to mention that we consider total expenditures and total revenues as ratios of GDP.

[9]These countries were chosen based on the availability of datasets and length of time series.

Table 2.1: Summary of empirical panel fiscal sustainability studies

Authors	Sustainability Test	Time 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
Afonso and Jalles (2012)	Cointegration between revenue and expenditure	OECD countries (1970 - 2010)	Fiscal stance sustainability not confirmed
Beqiraj et. al (2018)	Panel cointegration test between primary balance and public debt	21 OECD countries (1991 - 2015)	Fiscal stance judged to be unsustainable
Bökemeier and Stoian (2018)	Fiscal reaction function of primary balance and debt	CEECs (1997 - 2013)	Fiscal stance sustainable for selected countries
Brady and Magazino (2018)	Stationarity of public debt	19 European countries (1970 - 2016)	Fiscal stance sustainability confirmed
Chen (2014)	Stationarity of public debt	Selected European countries and G-7 group of countries ()	Stationarity of public debt confirmed for selected countries (allowing for nonlinearities)
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	Central and Eastern European countries (1990 - 2012)	Sustainable fiscal stance
Llorca and Redzepagic (2008)	Cointegration between revenue and expenditure	CEECs (1999:1 - 2006:1)	Fiscal stance sustainable in selected countries
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

Figure 2.2: Revenue and expenditure plot

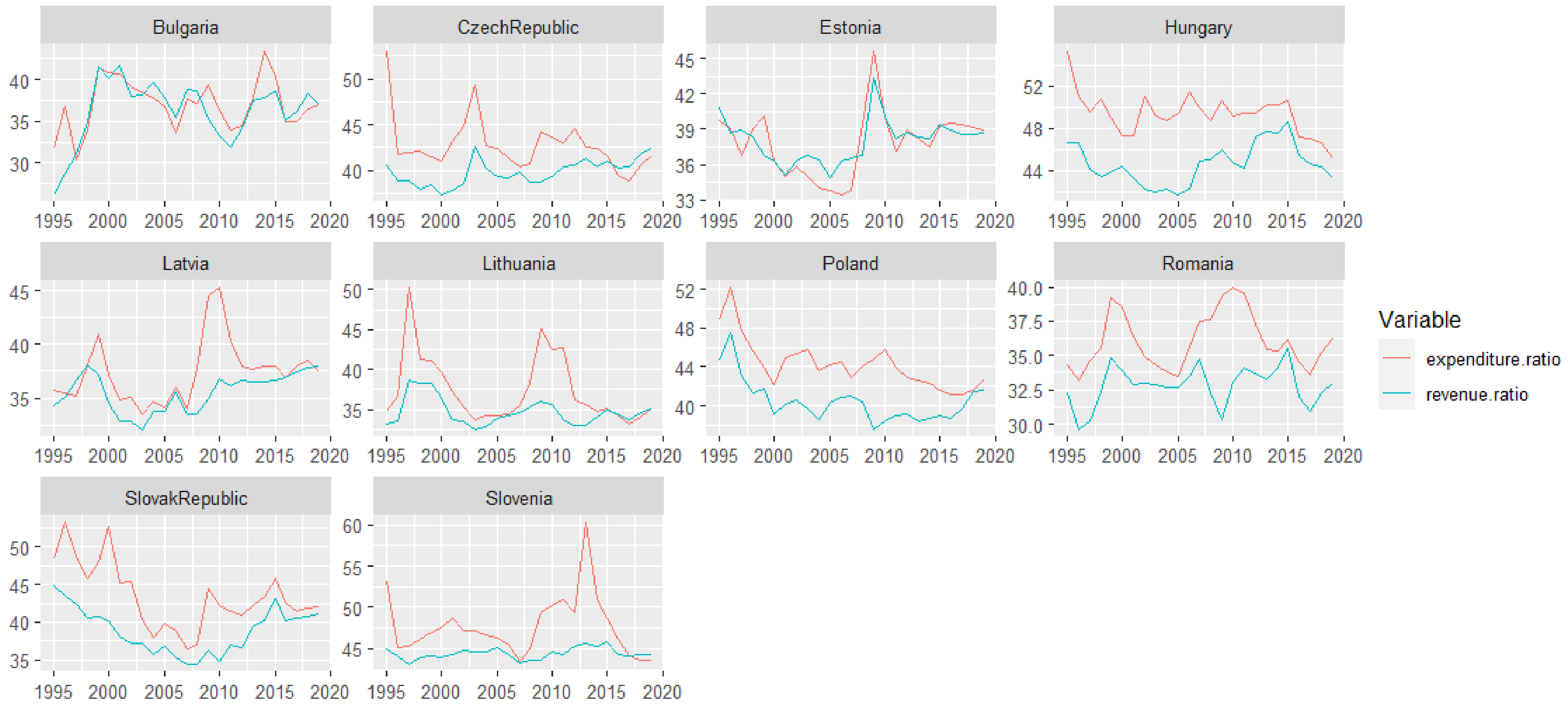


Figure 2.2 provides a graphical overview of revenue and expenditures for each country in the panel. We notice that in most of the cases, revenues and expenditure move in the same direction even though expenditures seems to be higher than revenues for most of the time periods. Poland, Hungary and Romania displayed high variability in the revenue-expenditure relationship. Since spending exceeds revenue in almost all cases, government has to find ways of financing the deficits either by borrowing or increasing taxes, for instance. An exceptional case is that of Estonia, where the variability between expenditure and revenue is low (revenue followed expenditure closely) even in the aftermath of the financial crisis. According to Friedrich and Reiljan (2015), the Estonian government addressed the crisis by implementing austerity-oriented fiscal policy measures in order to reduce inflation, reduce external trade positions, improve integration with the common market and overcome the banking crisis. Hence, they implemented measures to increase and stabilize revenues, such as increasing the tax burden (value added tax from 18% to 20%), reducing tax exemptions, increasing social security contributions among others. This explains the high revenues for Estonia, as can be seen in Figure 2.2. Figure 2.3 shows a scatter plot that reveals the relationship between revenues and expenditures with a linear fit. A positive upward-sloping relationship can be observed between the two variables, which provides some hints as to the nature of the relationship between the two fiscal variables.

The issue of structural break has received considerable attention in both theoretical and empirical econometric literature, notable among them includes Andrews et al. (1996), Andrews (1993), Bai and Perron (1998), among others. Structural breaks in the mean of data (or the changes in the coefficient of a regression), usually 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). They combined the F-statistics test by Andrews (1993) and Andrews and Ploberger (1994) to test the possibility of structural breaks and the technique by Bai and Perron (2003) to locate the break dates and optimal breaks in the individual series of the data.[10] Table 2.2 provides results of the break dates for both revenue and expenditures. Regarding revenues, the number of breaks ranges between 1 to 4. We notice that the majority of the breaks were recorded before the early 2000s, which could possibly represent a policy shift as most of the CEECs 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 countries, which could also be attributed to the exogenous shock and the consequences of the global financial crises. This provides justification for the presence of the breaks and the fact that it has to be accounted for in the modelling process.

[10]We select the optimal number of breaks by choosing the number of breaks with the least sum of square residuals.

Table 2.2: Dates of structural breaks for individual series

Countries (CEECs)	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

Figure 2.3: Panel scatter plot (with linear fit)

Table 2.3 provides a summary statistics 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 average, we observe that expenditures are higher than revenues, which is not so surprising since the role of government (spending) has become more important either to stimulate economic growth or in direct response to macroeconomic or external shocks. With respect to the individual countries, Hungary, Slovenia and to some extent, the Czech Republic have the highest revenue and expenditure to GDP ratios. These are relatively bigger economies in the CEEC.

Table 2.3: Summary statistics

	Expenditure to GDP ratio				***Revenue to GDP ratio***			
Country	mean	min	max	std dev	mean	min	max	std. dev.
Bulgaria	37.0	30.4	43.3	3.10	36.2	26.3	41.7	3.78
Czech Republic	42.7	38.9	53.0	2.96	39.8	37.4	42.5	1.39
Estonia	37.8	33.5	45.6	2.74	38.0	34.9	43.4	1.87
Hungary	49.4	45.1	55.2	2.0	44.6	41.7	48.7	1.93
Latvia	37.4	33.5	45.3	2.92	35.5	32.1	38.1	1.79
Lithuania	37.4	33.2	50.3	4.25	34.7	32.5	38.7	1.73
Poland	44.3	41.1	52.2	2.54	40.4	37.6	47.6	2.18
Romania	36.1	33.3	40.0	2.05	32.8	29.6	35.5	1.44
Slovakia	43.5	36.4	53.3	4.40	38.8	34.3	44.9	3.01
Slovenia	47.7	43.3	60.3	3.64	44.4	43.1	45.9	0.70

Next, we investigate the adherence of member countries to the SGP. As per the SGP requirements, member states of the EU are supposed 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 2.4 provides an overview of the fiscal balance (as GDP ratios) of the CEECs for the sample period where negative (positive) values indicate deficits (surplus). We notice that, with the exception of Estonia, which violated the SGP only once (1999), all other countries violated this rule a couple of times.The violations occurred mostly between 1995 and 1998, coinciding with the period prior to their EU accession. Secondly, during the financial crises and the European debt crisis periods between 2008 and 2012, we also noticed another round of SGP violations by the countries, with the exception of Estonia. CEECs have therefore, run fiscal deficits over the years and have not followed the 3% deficit limit rule strictly.

To begin with the empirical estimations, we test for stationarity of the variables. As part of the cointegration requirement, the variables should be integrated of order 1 ($I(1)$). In other words, if variables are not stationary at their levels, they must be stationary at their first difference. Firstly, we adopt the so-called first generational panel unit root test, which does not account for cross-sectional dependence in the panel. Notable among them are Levin, Lin and Chu (LLC) test and Im, Pesaran and Shin (IPS) unit root test. LLC test (see Levin et al. (2002)) is based on performing a separate Dickey-Fuller (DF) regression for each cross-section, estimating the ratio of the long-run to short-run standard deviations and computing the panel test statistics. The null hypothesis states that each individual time series contains a unit root against an alternative stationary series. The IPS unit root test (see Im et al. (2003))

Table 2.4: Fiscal balance

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 indicate violation of the EU Stability and Growth Pact. Source: author's computations

is similar to LLC, however, it allows for a heterogeneous coefficient of the lagged series and then averages the individual unit root test statistics. These unit root test (LLC and IPS) are limited in that they both assume cross-sectional independence, which may not be feasible, especially if one considers countries which are linked geographically and economically. The possibility of cross-country interdependence is high in such instances.

As a result, we adopt the cross-sectionally augmented Im, Pesaran and Shin test (CIPS), which accounts for CSD. CIPS unit root test approach entails augmenting the usual ADF regression with a lagged cross-sectional mean and its first difference to account for CSD. After running the regression for each unit, the test statistic of the lagged values is averaged to arrive at the CIPS statistics.

Tables 2.5 and 2.6 provides a report of the unit root test for revenue and expenditures. The null hypothesis of unit root for both LLC and IPS is strongly rejected at 1% significance level for both revenue to GDP ratio and expenditure to GDP ratio, irrespective of whether we consider a model with an intercept or trend. This implies that both variables are not integrated. However, if we consider a unit root test accounting for CSD (table 2.6), then the null hypothesis of a unit root cannot be rejected for both variables at their level whether we consider a model with an intercept or a trend (based on the p-values). Furthermore, we proceed by testing for stationarity at first difference taking cross-sectional dependence into consideration. The stationarity test for the first difference of the variables (table 2.6) shows the absence of a unit root. It is observed that there is enough evidence to reject the null hypothesis at 5% significance level when we consider a model with a constant for both revenue and expenditure. Considering a model with a trend, we observed that the null of unit root can be rejected at 5% level for revenue to GDP at the first difference. However, there is no compelling evidence of rejecting the null of unit root for the first difference expenditure with trend at 5% level (rejection at 10% significance level is possible). Based on the result and discussion, we therefore conclude that revenues and expenditure are stationary at their first difference and are $I(1)$.

Table 2.5: Panel unit root test - first generational test

	Revenue to GDP ratio				***Expenditure to GDP ratio***			
	LLC		***IPS***		***LLC***		***IPS***	
Country	test statistic	p-value	Test stat.	p-value	Test stat.	p-value	Test stat.	p-val.
Intercept	-5.088	0.000	-5.425	0.000	-6.819	0.000	-6.277	0.000
Trend	-4.243	0.000	-4.655	0.000	-4.7628	0.000	-4.710	0.000

Table 2.6: Panel unit root test - CIPS

	Revenue to GDP ratio				***Expenditure to GDP ratio***			
	Level		***First difference***		***Level***		***First difference***	
Country	statistic	p-value	Stat	p-value	statistic	p-value	Stat	p-val.
Intercept	-2.060	0.1	-2.777	0.01	-1.693	0.1	-2.823	0.01
Trend	-1.965	0.1	-2.945	0.040	-2.251	0.1	-2.794	0.081

2.3.1 Panel cointegration test

After establishing that revenue and expenditure are $I(1)$, we investigate 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 appealing because it accounts for structural breaks and cross-sectional dependence in panel data, making it very desirable. Secondly, the test is robust against serial correlation and heteroscedasticity in the residuals.

Suppose we consider y_{it}, which evolves according to :

$$y_{it} = \alpha_i + \theta_i t + \pi D_{it} + x_{it}' q_i + (D_{it} x_{it})' \mu_i + \epsilon_{it} \tag{2.10}$$

$$x_{it} = x_{it-1} + w_{it} \tag{2.11}$$

Where x_{it} is a k dimensional vector of regressors modelled as a random walk process, w_{it} is the error term (which follows a white noise process), $i = 1, ...N$ represents the cross sectional dimension, $t = 1, ...T$ represents the time period. D_{it} is a scalar representing a break dummy such that $D_{it} = 1$ if $t > T_i$ and zero otherwise. Parameters α_i and q_i are the intercept and slope respectively before the break, whilst π_i and μ_i are the changes in these parameters after the shift.

To allow for cross-sectional dependence, the residual ϵ_{it} is modelled using unobserved common factor as:

$$\epsilon_{it} = \lambda_i' C_t + v_{it} \tag{2.12}$$

$$C_{jt} = \rho_j C_{jt-1} + m_{jt} \tag{2.13}$$

$$\phi_i(L) \Delta v_{it} = \phi_i v_{it-1} + e_{it} \tag{2.14}$$

where $\phi_i(L) := 1 - \sum_{j=1}^{pi} \phi_{ij} L^j$ represents a scalar polynomial in the lag vector L, C_t represents a vector of unobservable common factor C_{jt} where $j = 1, ..., r$ and lastly λ_t represents vector of factor loading. To impose strict stationary, it is assumed that $\rho_j < 1$ for all j. This then implies from (2.12) that the order of integration of ϵ_{it} is dependent on the idiosyncratic disturbance term v_{it}. The data-generating relation in (2.10) is cointegrated provided $\phi < 0$ and spurious if $\phi = 0$. The null hypothesis of the test is that all N cross sectional units are spurious ($H_0 : N_1 = 0$) whilst the alternative hypothesis states that the first N_1 units are cointegrated with the remaining $N_0 := N - N_1$ considered spurious ($H_1 : N_1 > 0$).

Testing for the null as against the alternative hypothesis, Westerlund and Edgerton (2008) used a Lagrangian Multiplier (LM) principle which states that the score vector has a zero mean when estimated at the vector close to the true parameters under the null. They proposed two tests for the null hypothesis of no cointegration.

The proposed test is derived from a Lagrangian multiplier (LM) function in the similarity of Schmidt and Phillips (1992), Ahn (1993), and Amsler and Lee (1995).

The first test statistics is $Z_\tau(N)$ (based on t-ratio of the slope) whilst the second test statistic $Z_\phi(N)$ is based on the quotient between least square estimate of the residual slope and the standard deviation of the regression. Further detailed discussion about the test can be found in Westerlund and Edgerton (2008). The selection of optimum lag length is based on an automatic procedure adopted from Campbell and Perron (1991). We test the null of "no cointegration" against an alternative hypothesis of "cointegration" between revenue and expenditures.

Regarding the output of the test, three scenarios are considered. Firstly, we test the null hypothesis under the condition of absence of breaks and no factors. That is, we assume there are no breaks in the cointegration relationship and do not account for cross-sectional dependence. 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). A maximum of 2 breaks is chosen for the cointegration relationship. This can be justified, firstly, by the argument that most of these countries had to adjust their fiscal policy decision before joining the EU. Secondly, the external shock from the financial crisis affected the revenues and expenditures of these countries, leading to a potential structural break.

From Table 2.7, it can be noted that none of the processes is cointegrated when we consider significance at a strict 5% level. Considering a less strict significance level at 10%, we find evidence of cointegration for the model with no breaks (and no factors) for the $Z_\tau(N)$ test and no cointegration for $Z_\phi(N)$. Hence, even with no breaks, the cointegration relationship is not strongly confirmed.[11] The study presents evidence of the lack of cointegration between total revenues and expenditures (all GDP ratios). This implies a rejection of the fiscal sustainability hypothesis for CEECs which contrasts previous studies on CEECs notably by Krajewski et al. (2016) and Llora and Redzepagic (2007). Even though they both employed a panel cointegration procedure, their studies did not test (and account) for structural breaks and cross-sectional dependence in the cointegration relationships, which could be considered as a limitation. A model with no breaks from table 2.7 is comparable with their results. However, the results of the test is inconclusive since rejection of the null of no cointegration (at less strict 10% significance) only applies to $Z_\tau(N)$. Implementing a cointegration test that accounts for breaks and cross-sectional dependence in a panel data settings constitutes one of the main contributions of this study.

2.3.2 Adjusting fiscal variables for cyclicality

Recall from (2.8), which depicts the relationship between revenues and expenditures. We decompose these fiscal variables into trend and cyclical components. Following

[11] I would like to thank Joakim Westerlund for making the Gauss codes for the cointegration test available.

Table 2.7: Panel cointegration test of revenue and expenditure

	$Z_\tau(N)$		$Z_\phi(N)$	
Models	Value(τ)	P-Value	Value(ϕ)	P-value
No breaks (no factors)	-1.515	0.065*	-0.963	0.168
Level break	-0.983	0.163	-1.100	0.136
Regime shift	0.238	0.594	-0.200	0.421
Number of observations	250		250	

Westerlund and Edgerton (2008) cointegration test with maximum two breaks in the cointegration relationship. *, ** and *** denotes rejection of the null hypothesis at 10%, 5% and 1% respectively

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 could be influenced by business cycles or shocks. The trend component, on the other hand, represents an active discretionary fiscal policy and, hence, should be considered when examining the long-term behaviour of government policy. Decomposing (2.8), we get

$$R_t^c + R_t^\tau = \alpha + \gamma(GG_t^c + GG_t^\tau) + \mu_t \tag{2.15}$$

Where R_t^c and R_t^τ represent cyclical and trend components of revenue whilst GG_t^c and GG_t^τ 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 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 (2.15), we end up with

$$R_t^\tau = \alpha + \gamma(GG_t^\tau) + \mu_t \tag{2.16}$$

A popular tool for decomposing a series into trend and cyclical components is the HP filter (see Hodrick and Prescott (1997)). In a seminal paper, Hamilton (2018) proposed an alternative method for de-trending a series and showed 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. Finally, the values of

HP smoothing parameter are vastly at odds with common practice and therefore, not reliable. The study applied the alternative filter proposed by Hamilton to obtain the trend component of the revenue and expenditure series. Additionally, for the sake of the robustness of our results, we also applied the HP filter to obtain the trend component of the variables, after which the cointegration test is applied.

As a requirement for cointegration, we test the cyclically adjusted variables at their levels and first difference to ensure they are $I(1)$. Table 2.10 in the appendix indicates that cyclically adjusted expenditure and revenue have a unit root at their levels (considering significance at 5% level). However, they are stationary at the first difference if we consider a model with an intercept. We are therefore, able to proceed with the cointegration test.

Table 2.8 provides the result of the cointegration test (based on Westerlund and Edgerton (2008)) between cyclically adjusted revenues and expenditure. When we consider a model with no breaks and no account for cross-sectional dependence, we reject the null hypothesis at a strict 1% level, which implies there is enough evidence against the null hypothesis of no cointegration.

In contrast, the null of no cointegration cannot be rejected when we consider breaks in the level or a regime shift and account for cross-sectional dependence for both $Z_\tau(N)$ and $Z_\phi(N)$ tests (for the test designated "a", which allows for maximum two structural breaks). This implies that there is no evidence against the null and, hence, lack of support for cointegration. Furthermore, the result is mixed when we allow for one structural break (test designated "b"). The null of no cointegration is rejected when a model with a level break is considered and hence, evidence of cointegration. However, the null cannot be rejected for a regime shift model (breaks in the intercept and the trend) and therefore, no support for cointegration.

The above result does not provide support for cointegration of revenues and expenditures for the CEECs. This implies that if we consider cyclically adjusted variables, there is still not enough evidence that government in CEECs have jointly pursued a sustainable fiscal policy based on the cointegration analysis of their revenues and expenditures. Since cointegration has not been confirmed, estimating the cointegrating vector/slope is therefore, not very useful. As a robustness check, we test for cointegration with cyclically adjusted variables after extracting the trend using the HP filter (see table 2.9 in the appendix). The results are mixed depending on the number of structural breaks considered and hence, do not point to a more robust result. The evidence for cointegration is therefore, not strong enough to infer sustainability of fiscal stance for the CEECs.

2.4 Conclusion

This study sought to ascertain if the fiscal sustainability hypothesis holds for 10 CEECs from the period 1995 to 2019. Previous studies have shown that these coun-

Table 2.8: Panel cointegration test of cyclically adjusted revenue and expenditure

	$Z_\tau(N)$		$Z_\phi(N)$	
Models	Value(τ)	P-Value	Value(ϕ)	P-value
No breaks (no factors)	-2.842***	0.002	-3.587***	0.000
Level break[a]	0.157	0.562	0.247	0.597
Regime shift[a]	-0.328	0.371	-0.680	0.248
Level break[b]	-4.138***	0.000	-6.422***	0.000
Regime shift[b]	-1.018	0.154	0.324	0.627
Number of observations	220		220	

Westerlund and Edgerton (2008) cointegration test. Scenario "a" considers a maximum of two breaks, whereas in scenario "b" only one break is allowed in the cointegration relationship. The null hypothesis indicates "No cointegration". *, ** and *** denotes rejection of the null hypothesis at 10%, 5% and 1% respectively. The Hamiliton filter is used to obtain cyclically adjusted revenues and expenditures.

tries have pursued policies compatible with the government's intertemporal budget constraint. We tested the hypothesis of sustainability of the fiscal stance by examining the cointegration relationship between revenues and expenditures, both as ratios of GDP. The econometric intuition is that if revenues and expenditure can be expressed as a linear combination and their 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 CEECs. The result is also in sharp contrast to earlier panel studies conducted for CEECs, which have all pointed to the direction of cointegrated revenue and expenditures. However, none of the previous studies accounted for structural breaks and cross-sectional dependence in the cointegration regression, which is associated with macro panels. The study, therefore, tested, found evidence and accounted for structural breaks for CEECs, 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 and the subsequent European debt crises in 2009/2010.

As a next step, the study makes a justification for the use of 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 devoid of business cycle fluctuations or shocks. This is plausible because macroeconomic shocks usually induce an automatic response from policymakers and do not necessarily characterize discretionary policy. The results indicate limited evidence that cyclically adjusted revenue and expenditures are cointegrated. Considering the fact that these variables are ratios to GDP, a unit slope in the cointegration relationship is necessary to guarantee strong sustainability in the sense of Quintos(1995). Lack of evidence of cointegration implies that the debt-to-GDP ratio may not be bounded and therefore, not finite (debt-to-GDP will increase continuously, and the fiscal stance is judged not 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 debt. Secondly, the government may have difficulties in marketing its debts to new investors and therefore, would not be able to raise substantial additional revenue by issuing bonds in the future due to unattractiveness of its debts. Otherwise, the government would have to pay high interest to make its debt attractive to investors. CEEC governments may therefore, have to alter their fiscal policy by increasing revenue or reducing expenditure or both to counter the deficit problem. The study provides new evidence using cyclically adjusted revenue and expenditure for panel sustainability analysis in the context of CEECs. The discretionary action of the government is deemed insufficient to infer sustainability of the fiscal stance. The governments in CEECs must, therefore, do more to address the fiscal deficit problem by way of fiscal consolidation to avoid future implications of sustainability.

2.5 Appendix

Table 2.9: Panel cointegration test of cyclically adjusted revenue and expenditure (HP filter used for detrending series)

	$Z_\tau(N)$		$Z_\phi(N)$	
Models	Value(τ)	P-Value	Value(ϕ)	P-value
Level break[a]	-2.703**	0.003	-3.870***	0.006
Regime shift[a]	-0.141	0.444	-0.562	0.287
Level break[b]	0.519	0.698	-0.034	0.487
Regime shift[b]	-4.000***	0.000	-1.335*	0.091
Number of observations	250		250	

Westerlund and Edgerton (2008) cointegration test. In scenario "a" we considers a maximum of two breaks, whereas in scenario "b" only one break is allowed in the cointegration relationship. HP filter is used for de-trending of the series.

Table 2.10: Panel Unit root test for cyclically adjusted variables - CIPS

	Revenue to GDP ratio				***Expenditure to GDP ratio***			
	Level		***First difference***		***Level***		***First difference***	
Country	statistic	p-value	Stat	p-value	statistic	p-value	Stat	p-val.
Intercept	-2.322	0.060	-2.607	0.0135	-2.265	0.080	-2.882	0.01
Trend	-2.426	0.1	-2.369	0.1	-2.628	0.1	-2.726	0.1

Cross-sectional augmented Im, Pesaran and Shin (CIPS) unit root test. Null hypothesis implies non-stationarity. Maximum 2 lags are used for the test.

2.5.1 Proof of equation 2.6 by forward substitution

Substituting the expression $E_t = G_t + (\rho_t - \rho)B_{t-1}$ into 2.5 we get,

$$E_t + B_{t-1}(\rho - \rho_t + 1 + \rho_t) = R_t + B_t \tag{2.17}$$

$$E_t + B_{t-1}(\rho + 1) = R_t - B_t \tag{2.18}$$

Making B_{t-1} the subject from (2.18) ,

$$B_{t-1} = \frac{R_t}{(1+\rho)} - \frac{E_t}{(1+\rho)} + \frac{B_t}{(1+\rho)} \tag{2.19}$$

We can continue similarly with

$$B_{t-2} = \frac{R_{t-1}}{(1+\rho)} - \frac{E_{t-1}}{(1+\rho)} + \frac{B_{t-1}}{(1+\rho)} \tag{2.20}$$

Substituting 2.19 into 2.20 and re-arranging terms, we get

$$B_{t-2} = \frac{R_{t-1}}{(1+\rho)} + \frac{R_t}{(1+\rho)^2} - \Big(\frac{E_{t-1}}{(1+\rho)} + \frac{E_t}{(1+\rho)^2}\Big) + \frac{B_t}{(1+\rho)^2} \tag{2.21}$$

From 2.18, we can write the expression for B_{t-3} as

$$B_{t-3} = \frac{R_{t-2}}{(1+\rho)} - \frac{E_{t-2}}{(1+\rho)} + \frac{B_{t-2}}{(1+\rho)} \tag{2.22}$$

Substituting 2.21 into 2.22 and re-arranging terms, we get

$$\begin{aligned} B_{t-3} = \frac{R_{t-2}}{(1+\rho)} + \frac{R_{t-1}}{(1+\rho)^2} + \frac{R_t}{(1+\rho)^3} - \Big(\frac{E_{t-2}}{(1+\rho)} + \frac{E_{t-1}}{(1+\rho)^2} + \frac{E_t}{(1+\rho)^3}\Big) \\ + \frac{B_t}{(1+\rho)^3} \end{aligned} \tag{2.23}$$

We can continue to derive equation for B_{t-4} as per below

$$\begin{aligned} B_{t-4} = \frac{R_{t-3}}{(1+\rho)} + \frac{R_{t-2}}{(1+\rho)^2} + \frac{R_{t-1}}{(1+\rho)^3} + \frac{R_t}{(1+\rho)^4} - \Big(\frac{E_{t-3}}{(1+\rho)} + \frac{E_{t-2}}{(1+\rho)^2} \\ + \frac{E_{t-1}}{(1+\rho)^3} + \frac{E_t}{(1+\rho)^4}\Big) + \frac{B_t}{(1+\rho)^4} \end{aligned} \tag{2.24}$$

Doing this recursively up to B_{t-n} and summing from $s = t+1$ to some k, for all $s, k > 0$ we get

$$B_{t-n} = \sum_{s=t+1}^{k} \Big(\frac{1}{1+\rho}\Big)^{s-t} R_{s-n} - \sum_{s=t+1}^{k} \Big(\frac{1}{1+\rho}\Big)^{s-t} E_{s-n} + \Big(\frac{1}{1+\rho}\Big)^{s-t} B_{s-n} \tag{2.25}$$

Hence, the summation from $s = t+1$ to infinity, B_{t-1} can be written as

$$B_{t-1} = \sum_{s=t+1}^{\infty} \Big(\frac{1}{1+\rho}\Big)^{s-t} (R_{s-1} - E_{s-1}) + lim_{s\to\infty}\Big(\frac{1}{1+\rho}\Big)^{s-t} B_{s-1} \tag{2.26}$$

□

2.5.2 Data availability and scripts

Data used for this paper is available on my Github page (https://github.com/Benjamin-Owusu/Thesis). The script for the cointegration test was implemented using Gauss software and can be found on the page. The remaining script for the paper are implemented using R and can be found on the page. Accessibility to the data and scripts will be provided by the author upon reasonable request.

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Chapter 3

Assessing nonlinearities and heterogeneity in debt sustainability analysis: A panel spline approach

3.1 Introduction

The 2008/2010 European debt crisis revived the interest in public debt sustainability analyses. Further, the Covid-19 pandemic, which affected countries around the world, forced most governments to react in order to protect and cushion economies against the negative effects with the help of huge rescue and recovery programs that were mainly financed by issuing bonds. Thus, these policies affect both the current and future budget as well. Therefore, studying public debt sustainability and fiscal discretionary policies is as important as ever.

Based on the seminal contributions by Bohn (1995, 1998), we assess public debt sustainability by analyzing the reaction of the primary balance to changes in public debt relative to the gross domestic product (GDP). If a government reacts to a rise in the public debt ratio by actively adjusting its discretionary fiscal policy in terms of higher primary surpluses, the debt policy is considered to be sustainable. Usually, the relationship is empirically tested in a simple linear regression model. However, standard econometric specifications, assuming linearity between response and explanatory variable, could lead to model misspecification if the true data-generating process is nonlinear and, hence, could lead to wrong inferences. Therefore, it seems necessary to resort to non-parametric or semi-parametric models.

The need to increase the size of the dataset to improve statistical inference and to study the dynamic relationships between variables has made panel studies popular in applied macroeconomics research. This paper intends to get additional insights by going beyond the standard linear fixed effects model used in panel debt sustainability analyses. Thus, we do not only assume that the response of the discretionary fiscal policy is uniform across the distribution of the debt ratios, but we allow it to vary with the magnitude of the debt ratios. Particularly, we can show that the reaction of the structural primary balance changes with the size of the debt ratio, implying that in situations with low debt ratios, the relationship between the primary balance and debt ratio is expected to be different, compared to situations with high debt ratios. We consider two clusters of EU economies to allow for country specifics and distinguish between economic crisis and normal times, where we use AMECO data for 25 EU economies from 2000 to 2019. The reasoning for focusing on this sample is against the background of the sovereign debt crisis in Europe around the year 2010. The estimations demonstrate that the reaction to higher debt ratios differs in the two clusters and is higher in times of crisis than in normal times. Thus, our estimations support the hypothesis of heterogeneity in the data. These refinements are particularly important for policy implications and help to improve recommendations as they indicate that the status of the current debt and the economic situation is essential for the assessment and evaluation of fiscal sustainability. Our results show that, yes, size does matter. The level of the explanatory variable influences the reaction coefficient and, thus, sustainability. This means policy recommendations need to consider the status of the current debt situation in order to be successful, as the reactions show different behavior for low debt levels compared to medium and high ratios.

The contribution of the paper is threefold. Firstly, we applied a non-parametric model (panel penalized splines) to study the nonlinearities in debt sustainability analysis for EU countries. Hitherto, only a limited number of papers had explored nonlinear models for debt sustainability analysis in the panel data context. The panel penalized spline models provide the opportunity to visualize the behaviour of the reaction coefficient over the distribution of the covariate of interest (lagged debt ratio). Secondly, the paper employs a clustering algorithm to segregate the datasets due to the potential heterogeneity of the countries in the panel. This enabled us to study the debt sustainability characteristics of each cluster in the EU. Finally, the paper adds to the empirical literature of panel nonlinear models in debt sustainability analysis context.

The rest of this paper is organized as follows. The next section 3.2 presents a brief literature overview. Section 3.3 briefly discusses the theoretical background, and section 3.4 presents the estimation method and the outcomes. Finally, section 3.5 summarizes the main results and concludes.

3.2 Literature overview

Statistical testing of whether the inter-temporal budget constraint of a government is fulfilled began with the paper by Hamilton and Flavin (1986). Those authors tested whether the present-value borrowing constraint of the US federal government holds with annual data from 1960 to 1984. They did so by analyzing if the time series of the US federal public debt contains a bubble term that would indicate that public debt exceeded the present value of expected future primary surpluses, implying an unsustainable debt policy. Hamilton and Flavin performed several tests and found evidence of sustainability of the US federal debt policy.

The test procedure applied by Hamilton and Flavin has been criticized by Wilcox (1989) because their test does not allow for a stochastic interest rate. Therefore, he proposed a test where the discounted time series of public debt should be analyzed, and if that series converges to zero, sustainability of the public debt would be given. Applying that test to the same time series Hamilton and Flavin used, Wilcox finds evidence that the US federal debt is unsustainable.

The result obtained by applying the test proposed by Wilcox depends on the interest rate with which the series of public debt is discounted. Since this is a random variable and realizations of that variable in the past do not give information about future interest rates, it has been argued that tests should be resorted to that yield outcomes which are independent of the interest rate. Hakkio and Rush (1991), for example, tested for the cointegration of revenues and expenditures. The idea behind that approach is that a cointegrating relation between spending and revenues implies a stationary first difference and, thus, a sustainable debt policy for a positive interest rate.

Another test that does not rely on the interest rate is the one proposed by Trehan and Walsh (1991). They test whether public deficits, inclusive of interest payments, grow at most linearly. If that property is fulfilled, a given series of public debt is sustainable because any time series that grows linearly converges to zero if it is exponentially discounted, provided the real interest rate is positive. Denoting by B public debt (nominal terms) and by r the interest rate, another test proposed by Trehan and Walsh (1991) is to analyze whether a quasi-difference of public debt, $B_t - \lambda B_{t-1}$ with $0 \leq \lambda < 1 + r$, is stationary and whether public debt and primary surpluses are cointegrated. If government debt is quasi-difference stationary and public debt and primary surpluses are cointegrated, public debt is sustainable.

The dependence of sustainability tests on the interest rate has been heavily criticized by Bohn (1995, 1998) since future realizations of that random variable are unknown. Therefore, he tested whether the primary surplus relative to GDP is a positive function of the debt-to-GDP ratio that rises at least linearly with higher debt-to-GDP ratios. The intuition behind that procedure is that a positive reaction of the primary surplus to higher debt relative to GDP, implies mean reversion of the debt-to-GDP ratio and a stationary debt-to-GDP ratio is sustainable in a growing economy. This test has a plausible economic intuition: if governments run into debt today, they have to take corrective actions in the future by increasing the primary surplus. Otherwise, public debt will not be sustainable. Many applications of such a fiscal response function as suggested by Bohn, have followed. A recent overview can be found, for instance, in Beqiraj et al. (2018). Some applications have focused on nonlinear fiscal behaviour and studied "fiscal fatigue", which is a reverse in behavior of the primary balance as debt ratios become very high, the response peters out and becomes negative. This was introduced by Gosh et al. (2013) and studied by Checherita-Westphal and Zdarek (2017) as well as Fournier and Fall (2015) among others. Table 3.1 below provides a summary of papers regarding European debt sustainability in the panel data context.

Table 3.1: Summary of panel fiscal sustainability research in the EU context

Authors	Sustainability Test	Period and Country	Findings
Afonso et. al (2019)	Fiscal reaction function	28 EU countries (1970 - 2015)	Sustainable fiscal stance
Baldi & Staehr (2016)	Fiscal reaction function of primary balance, debt and business cycle variables	Different groups of EU countries (2001 - 2004)	Sustainable fiscal stance for all groups post-financial crisis
Ballabriga & Martinez-Mongay (2005)	Fiscal reaction function	15 EU countries (1977 - 2002)	Sustainable fiscal stance
Checherita-Westphal & Zdarek (2017)	Fiscal reaction function of primary balance response to debt	18 Euro Area countries (1970 - 2013)	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
Medeiros (2012)	Fiscal reaction function and VAR	15 EU Member States	Fiscal stance sustainability in the weaker sense
Schalk (2012)	Fiscal reaction function	12 Eurozone Countries (1999 - 2010)	Varied fiscal response depending on the quantile.
Stoian et al. (2022)	Fiscal reaction function	26 European Union countries (2005 - 2018)	Varied fiscal response depending on the quantile
Weichenrieder and Zimmer (2014)	Fiscal reaction function	Euro area countries (1970 - 2011)	Sustainable fiscal stance

3.3 Theoretical Background

To get a deeper insight into the logic of the test proposed by Bohn (1995, 1998) and the fiscal response function, we consider the accounting identity describing the accumulation of public debt in continuous time that is given by the following differential equation:

$$\dot{B}(t) = r(t)B(t) - S(t), \tag{3.1}$$

with $B(t) > 0$, the real public debt at time t, $r(t) > 0$ is the real interest rate and $S(t) \in \mathbb{R}$ is the real government surplus exclusive of interest payments. A government is said to follow a sustainable debt policy if its inter-temporal budget constraint is fulfilled, i.e. if the present value of public debt converges to zero asymptotically. The latter means that $\lim_{t\to\infty} e^{-C_1(t)}B(t) = 0$ holds, with $C_1(t) = \int_{t_0}^{t} r(\mu)d\mu$ the discount rate.

Assuming that the government in the economy determines the primary surplus to GDP ratio, $s(t) = S(t)/Y(t)$, such that it is a positive linear function of the debt-to-GDP ratio, $b(t) = B(t)/Y(t)$, and of a term that is independent of public debt, $\phi(t)$ (see Bohn (1995, 1998), Greiner and Fincke (2015), and Afonso and Jalles (2019)), the fiscal reaction function can be written as

$$s(t) = \psi\, b(t) + \phi(t), \tag{3.2}$$

where ψ is the reaction or response coefficient determining the change of the primary surplus to variation in public debt, relative to GDP. The parameter $\phi(t) \in \mathbb{R}$ is affected by other economic variables, such as social spending or transitory government expenditures in general. As regards $\phi(t)$, we posit that it is bounded from above and from below by a certain finite number that is constant over time. We should also like to emphasize that $\phi(t)$ cannot be completely controlled by the government. The government can influence that coefficient to a certain degree, but it does not have complete control over it because $\phi(t)$ is also affected by the business cycle (for instance), that can affect temporary government outlays.

Using (3.1) and (3.2), the period budget constraint of the government is obtained as

$$\dot{B}(t) = (r(t) - \psi)\, B(t) - \phi(t)\, Y(t). \tag{3.3}$$

Integrating equation (3.3) and multiplying both sides by $e^{-C_1(t)}$ leads to

$$e^{-C_1(t)}B(t) = e^{-C_3(t)}\, B(t_0) - e^{-C_3(t)} \int_{t_0}^{t} e^{-C_1(\mu)+C_2(\mu)+C_3(\mu)}Y(t_0)\phi(\mu)d\mu, \tag{3.4}$$

with $g(t)$ the growth rate of GDP and $C_1(\mu) = \int_{t_0}^{\mu} r(\nu)d\nu$, $C_2(\mu) = \int_{t_0}^{\mu} g(\nu)d\nu$, $C_3(\mu) = \psi\mu$. As regards the interest rate, we posit that the interest rate on government bonds exceeds the growth rate of GDP on average, $\int r(\mu)d\mu > \int g(\mu)d\mu$. We make this assumption because otherwise, the inter-temporal budget constraint would not pose a problem for the government since it can grow out of debt in that case.

Now, assume that ψ is positive on average so that $\lim_{t\to\infty} C_3(t) = \infty$ holds. Then, we get $\lim_{t\to\infty} e^{-C_3(t)} B(t_0) = 0$, and the first term in equation (3.4) converges to zero.

Since $|\phi(t)| < \infty$, we can set $\phi(t)Y(t_0) = 1$ so that it is sufficient to consider for the second term on the right-hand side in (3.4) the following expression,

$$\Omega(t) = \frac{\int_{t_0}^{t} e^{-C_1(\mu)+C_2(\mu)+C_3(\mu)} d\mu}{e^{C_3(t)}}$$

If $\int_0^\infty e^{-C_1(\mu)+C_2(\mu)+C_3(\mu)} d\mu$ remains bounded $\lim_{t\to\infty} C_3(t) = \infty$ guarantees that Ω converges to zero. In case of $\lim_{t\to\infty} \int_{t_0}^{t} e^{-C_1(\mu)+C_2(\mu)+C_3(\mu)}\ d\mu = \infty$, the limit of Ω is obtained by applying l'Hôpital as

$$\lim_{t\to\infty} \Omega(t) = \lim_{t\to\infty} \frac{e^{-C_1(t)+C_2(t)}}{\psi}$$

These considerations demonstrate that a strictly positive reaction coefficient on average $\psi > 0$ and a positive interest rate-growth rate gap $\lim_{t\to\infty} \int r(\mu)d\mu - \int g(\mu)d\mu = \infty$ must hold to imply that the debt policy of a government is sustainable (see Greiner and Fincke (2015), chap. 2.2-2.5 for details). The latter requirement means that the interest rate-growth rate difference must be positive on average in order for the present value of public debt to converge to zero asymptotically. It must be pointed out that the reaction coefficient ψ can be varying, and it may even be negative for some time periods. However, on average, that coefficient must be positive. The aspect regarding the interest rate-growth rate interval becomes particularly important when considering ratios to GDP. A shortcoming of the former analysis is that it implicitly assumes that the primary surplus can grow without an upper bound. However, a positive but small reaction coefficient on average does not necessarily guarantee a bounded debt-to-GDP ratio. It can be demonstrated that the public debt-to-GDP ratio $b(t)$ remains bounded if the reaction coefficient $\psi > \int_{t_0}^{t} (r(\mu) - g(\mu))\, d\mu$, while it diverges to infinity in the case of $\psi \le \int_{t_0}^{t} (r(\mu) - g(\mu))\, d\mu$. Thus, the reaction coefficient must exceed the difference between the interest rate on public debt and the GDP growth rate on average, such that the debt-to-GDP ratio remains bounded.

Here, we should like to stress three aspects. Firstly, the previous calculations show the significance of the difference $\int r(\mu)d\mu - \int g(\mu)d\mu$. If the GDP growth rate exceeds the interest rate on public debt, the debt-to-GDP ratio remains bounded. In that case, the government can grow out of debt, as already mentioned above. Secondly, a positive reaction coefficient that falls short of the difference between the interest rate and the GDP growth rate implies a rising debt-to-GDP ratio if the interest rate exceeds the GDP growth rate. But, such a policy is not sustainable because it would require permanently rising primary surplus to GDP ratios, which is not feasible because the primary surplus relative to GDP is bounded from above and cannot exceed the GDP. Thus, there exists a critical threshold value of the debt ratio beyond which sustainability is excluded. In that case, the outstanding debt ratio is too large to be covered by the sum of discounted future primary surpluses (cf. Greiner and Fincke, 2015, chap. 2.1, Prop. 3.).

To finish our theoretical considerations, we want to point out that the reaction coefficient could be varying. Times series analyses provide empirical evidence that this coefficient is not constant, for instance, over time (see Greiner and Fincke, 2015, chap. 2.2-2.5). We emphasize this point because in the empirical part, we estimate a nonlinear model describing the relationship between the primary surplus and public debt.

The use of a nonlinear model is of interest because the question arises of which factors are responsible for variations in the coefficients. When analyzing the response of the primary surplus to variations in public debt, there is evidence that the reaction depends on the magnitude of the public debt ratio. For example, Ghosh et al. (2013) analyzed 23 advanced economies from 1970-2007. They found that the reaction decreases once a critical value has been passed (what they refer to as 'fiscal fatigue'). Checherita-Westphal and Zdarek (2017) also explored 'fiscal fatigue' in euro-area economies with a focus on the primary balance.

3.4 Empirics

3.4.1 Estimation method

The empirical estimation methodology uses parametric (linear) and non-parametric econometric models. For the nonlinear model (non-parametric), we resort to the penalized splines fixed effects estimator, according to Puetz and Kneib (2018). Such additive non-parametric or semi-parametric models have become increasingly popular in empirical works (Baltagi and Li, 2002). We argue that this approach is desirable because non-parametric modeling does not impose restrictions regarding the functional relationship between the response variables and the regressors. Rather, the functional shape of the co-variate effect is derived from the datasets. Pioneering works on penalized splines can be traced back to Hastie and Tibshirani (1990) with their introduction of generalized additive models, which provide a flexible way of modelling the response of a dependent variate to the co-variates. Subsequent contributions were made by Wood (2000), who introduced the mixed model representation of penalized splines, as well as Ruppert et al. (2003), Eilers and Max (1996) and Kauermann (2005) among others.

In panel data settings, the correlation between covariates and the unobserved time-invariant factors in the error term is prevalent. Baltagi and Li (2002), Su and Ullah (2006) and Henderson et al. (2008) have all proposed alternative semi-parametric fixed effects estimators where the nonlinearity is modeled via a kernel estimator. However, Puetz and Kneib (2018) argued that modelling fixed effects and nonlinearity with the Kernel estimator is computationally demanding, especially with large datasets. Penalized splines, on the other hand provide a flexible and convenient way of modeling the nonlinearity without complications, and the fixed effects in the model can be conveniently handled by way of a first difference operator.

Following Puetz and Kneib (2018), we consider the following panel additive model

$$y_{i,t} = \mu_i + \sum_{g=1}^{p} f_g(x_{gi,t}) + v_{i,t} \tag{3.5}$$

where $i = 1, ..., N$ represents the individual countries and $t = 1, ...T$ represents the time period. $y_{i,t}$ is the response variable, whilst μ_i accounts for the time-invariant unobserved heterogeneity also known as the individual fixed effects in this case. We allow the unobserved heterogeneity to correlate with the regressors instead of the error term. The variable $v_{i,t}$ represents the error term, which is assumed to be normally distributed with zero mean and a constant variance. The functions $f_1(x_{gi,t})$,..., $f_p(x_{gi,t})$ are the nonlinear effects of the p continuous co-variates which can be approximated by a penalized B-splines according to Eilers and Marx (1996) together with a penalty term applied to the penalized least squares criterion. The aim of the penalty term is to penalize too much variability in the model as a way of regularizing the estimation in order to deal with overfitting to the data.

Assuming that (3.5) holds for each point, we could express the lag as

$$y_{i,t-1} = \mu_i + \sum_{g=1}^{p} f_g(x_{gi,t-1}) + v_{i,t-1} \tag{3.6}$$

in order to eliminate the fixed effects, we subtract (3.6) from (3.5) to obtain

$$\Delta y_{i,t} = \sum_{g=1}^{p} [f_g(x_{gi,t}) - f_g(x_{gi,t-1})] + v_{i,t} - v_{i,t-1} \tag{3.7}$$

where Δ is the difference operator. The nonlinear function f_g is approximated by the weighted sum of d_g B-spline basis functions $B_{g1}, ..., B_{gd}$ such that

$$f_g(x_{gi}) = \sum_{j=1}^{d_g} B_{gj}(x_{gi})\beta_{gj} = z_g^{'}(x_{gi})\beta_g. \tag{3.8}$$

Inserting (3.8) into (3.7) and applying the difference operator once again, we have

$$\Delta y_{i,t} = \sum_{g=1}^{p} [\Delta z_g(x_{gi,t})]^{'} \beta_g + \Delta v_{i,t}. \tag{3.9}$$

For the sake of simplicity, (3.9) can be written in compact matrix notation as

$$\Delta y = \sum_{g=1}^{p} \Delta z_g \beta_g + \Delta v \tag{3.10}$$

where $\Delta y = (y_{12} - y_{11}, ..., y_{N,T} - y_{N,T-1})'$ is a column vector with dimension $N(T-1)$ and Δz_g is derived by taking the difference between z_g and its lags which is of dimension $N(T-1)$ multiplied by d_g.

To obtain a first difference penalized spline estimator, the penalized least square criterion below is minimized

$$[\Delta y - \sum_{g=1}^{p} (\Delta z_g)\beta_g]'[\Delta y - \sum_{g=1}^{p} (\Delta z_g)\beta_g] + \sum_{g=1}^{p} \lambda_g \beta_g^{'} k_g \beta_g \tag{3.11}$$

where the matrix k is introduced and assigned to each smooth function to penalize too much variability of the adjacent coefficient in the vector β_g. This prevents overfitting the model to the data as mentioned earlier. λ_g is the smoothing parameter, which is the weight placed on the penalty term in the minimization criterion in (3.11). In reality, the smoothing parameters are unknown ex-ante. However, they can be estimated via maximum likelihood estimation by making use of the mixed model representation of the penalized splines (Puetz and Kneib (2018)). To estimate the smooth function f_g, equation (3.5) is identified such that $\sum_{i=1}^{N} \sum_{t=1}^{T} f_g(x_{gi,t}) \approx z_g \beta_g = 0$, for all $g = 1, \ldots, p$. In order to account for serial correlation in the residuals, which could arise as a result of first difference transformation, a Generalized Least Square (GLS) approach is used where the differenced model matrix (Δz_g) and Δy in (3.11) are multiplied by a block diagonal matrix. This ensures that the estimator is robust against serial correlation resulting from first differencing. Another appealing feature of the semi-parametric fixed effects estimator is the estimation of the derivative of the smooth function as well as the computation of simultaneous confidence bands for inferences. See Wiesenfarth et al. (2012) and Puetz and Kneib (2018) for details regarding the simultaneous confidence bands.

3.4.2 Model specification

In this paper, we model a fiscal reaction function in analogy to Bohn (1998), where the primary balance is assumed to be a linear function of debt relative to GDP and of other macroeconomic variables which serve as control variables. To obtain both linear and nonlinear estimates of the reaction function, we first model the primary balance as a linear reaction function of lagged debt and other macroeconomic variables as below:

$$S_{i,t} = \mu_i + \alpha X_{i,t-1} + \sum_{j=1}^{m} \beta_j Y_{it}^{j} + v_{i,t} \tag{3.12}$$

Where S is the response variable which in this case is the primary balance, X is the covariate of interest (lagged debt), and Y represents the set of control variables and $j = 1, ..., m$ refers to the number of control variables. α and β measures the impact of the linear regressors and control variables on the response variable respectively, and v is the uncorrelated error term assumed to be centred around zero with a constant variance. i depicts the individual countries in the panel data whilst t corresponds to the time series dimension.

In the second specification (in contrast to Bohn's linear specification), we do not posit that the response of the primary surplus to public debt is linear, but we estimate a non-parametric model of the form

$$S_{i,t} = \mu_i + f(X_{i,t-1}) + \sum_{j=1}^{m} f(Y_{i,t}^{j}) + v_{i,t} \tag{3.13}$$

where $S_{i,t}$ is the response variable, $X_{i,t-1}$ and $Y_{i,t}$ represent the covariate of interest and controls variables respectively. They both enter the model nonlinearly and hence represent the penalized functions. f is the nonlinear function which is approximated by penalized B-splines. $v_{i,t}$ is the uncorrelated error term assumed to be centred around zero with a constant variance. We shall apply both (3.12) and (3.13) to estimate the fiscal reaction function.

As a first step, we estimate the linear part of the model after which the nonlinear model is estimated. Regarding the nonlinear specification (3.13), all variables are modelled non-parametrically, meaning that we do not impose any restrictions on the relationship between the response variables and the covariates. For instance, if the true relationship between any of the covariates and the response variable is indeed linear, the spline model will impose a linear relationship via the effective degree of freedom of the co-variate in question. The effective degree of freedom is a meaningful and scale-free measure of the complexity of the penalized spline fit (Harezlak et al., 2018). This approach is particularly advantageous since we thus, avoid misspecification of the model that would result if apriori restrictions were imposed that are inconsistent with the true data-generating process.

The source of the data used for the empirical study is the European commission AMECO website (AMECO, 2021). Regarding the response variable, we use the cyclically adjusted primary balance, which represents the primary balance if the economy is at the potential level of aggregate production (adjusted for cyclicallity). In general, the primary balance is the fiscal balance of the government excluding interest payments. We focus on the cyclically adjusted balance because it is devoid of shocks or one-off fluctuations and because this is consistent with the EU fiscal framework (see Mourre et al., 2013). It is expressed relative to GDP, with positive (negative) values indicating surpluses (deficits). The explanatory variables include the lagged debt ratio of the previous period $t - 1$, which constitutes our main covariate of interest. Following the Barro (1979) tax smoothing principle[1], we use the business cycle variable (YVAR) and the public expenditure gap (GVAR) as control variables. YVAR, also known as the output gap is computed as the deviation of output (GDP) from its long-term trend. Similarly, we compute the GVAR as the deviation of real government spending from its long term trend. We use the HP filter to estimate the trend components of output and real spending. In order to capture the influence of international trade, we include the net export-ratio, that is given by the difference between exports and imports as a ratio to GDP. In addition, we include inflation in order to explore its influence on the primary balance as a proxy for monetary policy effects.[2] A total of 25 EU countries constitutes the sample with the exception of Lithuania and Croatia due to unavailability or missing data. The sample period is from 2000 to 2019. Hence a total of 500 observations in annual frequency were generated.

Before proceeding with the modelling process, we present a correlation matrix to show the relationship between all variables. Table 3.17 in the appendix depicts a positive relationship between the primary balance and lagged debt ratio even though the correlation coefficient is not so strong. Inflation and output gap all have negative relationships with primary balance. We also show the behaviour of the primary balance and lagged debt-ratio in figure 3.9 and figure 3.8 respectively (in the appendix) for selected countries. One can observe a sharp decline in the primary balance ratio during the 2009/2010 period for the bigger economies notably Germany, France, Portugal, Spain, Poland and the Netherlands. Regarding the debt ratio variable, one can observe an upward trend after 2008 for most of the countries. This supports the narrative that fiscal policy for most EU economies altered after the financial/European debt crisis period towards a more counter-cyclical approach.

3.4.3 Panel data econometric tests

3.4.3.1 Individual and time effects

To proceed with the data modelling, it is important to ascertain whether the individual and time dimensions in the panel datasets are significant. In other words, we test if either the individual, time or both effects should be considered in the panel data specification. We resort to the F-test for the presence of the said effects, where the nested Ordinary Least squares (OLS) model is compared to the "within estimator". The null hypothesis implies the absence of effects (be it individual, time or both), and the statistics is distributed as an F-distribution with the degree of freedom equal to the number of individuals less one for the numerator and the degree of freedom of the within model for the denominator (Croissant and Millo, 2019).

[1]According to the tax smoothing principle, public finance should be designed such that the tax rate is constant over time to minimize the excess burden of taxation.

[2]We used changes in the GDP price deflator as a proxy for inflation, with 2015 as the base year.

Table 3.2 provides a report of the F-test for the presence of time effect, individual and both effects with their respective test statistics, probability values and degree of freedom. It can be seen that for the time effects, the p-value suggests the null hypothesis cannot be rejected, and hence indicates the non-significance of only time effects. Individual effects, however, are significant (based on the p-value), and finally, a model with both individual and time effects is significant since the p-value indicates rejection of the null. Hence, it is feasible to estimate a panel model with either individual effect or both individual and time effect. A model with only time effect is not feasible, as can be seen from the test results.

We proceed to estimate a model with individual fixed effects as the main model specifications. Secondly, we provide estimates for a model with both individual and time fixed effect as a robustness check and show that the results do not differ significantly. Regarding the panel splines, one can augment the model with time (yearly) dummies as a way of accounting for time effects.

Table 3.2: Testing for individual and time effects - full sample

	Time effects			***Individual effects***			***Individual and time effects***		
Variables	Test-stats.	P-value	Df1 / Df2	Test-stats	P-value	Df1 / Df2	Test-stats	P-value	Df1 / Df2
Values	1.217	0.239	19 / 475	9.434	0.000	24 / 470	6.040	0.000	43 / 451
Num Obs	500			500			500		

F-test for individual and or time Effects. The null hypothesis indicates the absence of significant effects (be it individual, time or both)

Table 3.3: Panel poolability test - full sample

	Pooled OlS and Pvcm				***Fixed effects and Pvcm***			
Variables	F-stats.	P-value	Df1	Df2	F-stats	P-value	Df1	Df2
Values	16.70	0.000	144	350	11.17	0.000	120	350
Num Obs	500				500			

Chow test for poolability of panel data. Fixed effects "within model" and "pooled panel model" are compared to Panel Variable Coefficient Model (pvcm). The null hypothesis indicates model stability and hence both models are comparable. The alternative hypothesis implies model instability. The estimated model is as follows $S_{i,t} = \mu_i + \alpha X_{i,t-1} + \sum_{j=1}^{m} \beta_j Y_{it}^{j} + v_{i,t}$, where the response variable represents the primary balance ratio, the right-hand side variables includes the lagged debt-to-GDP ratio (X_{it}) and macroeconomic control variables, notably output gap, inflation and trade, which are all represented by vector notation Y_{it}^{j}.

3.4.3.2 Pooling of parameter estimates (slope homogeneity)

Before we begin with the estimations, we ascertain if the parameters in the panel specification can be pooled across the cross-sectional unit for the countries. A significant question that could arise is whether the parameters of the fiscal reaction function equation vary over time or change across the countries. Baltagi (2021) argues that one of the main motivations for pooling has to do with widening the data in order to obtain better and more reliable estimates of the model parameters. However, imposing stringent restrictions may come at some risk because the model will be biased if the restrictions are false. The decision to pool or not to pool is therefore, an important step in panel data econometrics. Regarding the pooling hypothesis, the idea is to compare a restricted model consisting of constant parameters over the cross-sections to an unrestricted model with variable parameters. We apply the Chow test, which is based on the F-test of stability for the panel coefficients, under the assumption that the residuals of the error term are normally distributed and homoskedastic (see Baltagi (2021)). Firstly, we compare a pooled panel linear model with an identical intercept and slope coefficient to a panel variable coefficient model (variable intercept and slope coefficient). Secondly, we compare a model with fixed effects "within-estimator" (variable intercept but identical slope coefficient), to a panel variable coefficient model. The results displayed in table 3.3 reveal a rejection of the null hypothesis of model stability for both models, which implies that the model parameters cannot be pooled since neither pooled OLS nor fixed effects models are comparable to a panel variable coefficient model. Hence, it is not always prudent to assume slope or coefficient homogeneity for the panel dataset. This may be due to high degree of heterogeneity in the data or a potential structural break in the data-generating process.

3.4.4 Clusters

Due to the data heterogeneity, we explore the possibility of dividing the countries into clusters where each cluster is made up of countries with similar characteristics. This is done to aid pooling the panel data since countries with similar characteristics will be categorized in the same sub-sample. We employ the K-means approach (see Hartigan and Wong (1979)), which entails finding clustres and cluster centers in the data set. Using this clustering technique, we tentatively choose the number of desired cluster centres and allow the K-mean algorithm to iteratively move the centers in a way as to minimize the sum within the variance of the cluster. Alternatively, an optimal number of clusters can be estimated by choosing the number of clusters for which the total within-cluster sum of squares is minimized. We estimate the clusters based on our variables of interest, notably the primary balance and the lagged debt ratio. Firstly, we use the raw variables, after which standardized variables are used for the clustering. The results of the two cluster characteristics can be found in Table 3.4.[3]

Table 3.4: K-mean summary statistics

	Non-standardized variables		***Standardized variables***	
Clusters	lagged debt	Prim. balance	lagged debt	Prim. balance
Cluster 1	104.6%	0.7%	132.6%	36.1%
Cluster 2	42.2%	-0.25%	-47.1%	-12.8%

With reference to the non-standardized variables, Cluster 1 has an average debt-to-GDP ratio of 104% with positive primary balance, whilst Cluster 2 has a debt-to-GDP ratio of 42% with negative primary balance. This is intuitive since a higher debt-to-GDP ratio requires that a higher primary balance is generated to ensure the debt is repayable. Results from the standardized variable are not intuitive since the variables have been transformed so they lose their original meaning. However, using the standardized variables for clustering ensures that the variables used for clustering have comparable units, and hence, we avoid scaling problems. A graphical view of the cluster plots are shown in figure 3.7 in the appendix. Cluster 1 is made up of the following EU countries: Belgium, Greece, France, Italy, Cyprus, Austria and Portugal. Countries that make up Cluster 2 are as follows: Bulgaria, Czechia, Denmark, Estonia, Germany, Ireland, Spain, Latvia, Luxembourg, Hungary, Malta, Netherlands, Poland, Romania, Slovenia, Slovakia, Finland and Sweden. Since there are 20 data points for each country, we classify a country into a particular cluster if more than half of the data points for that country belong to that specific cluster. For instance, Bulgaria has two data points in Cluster 1 and 18 data points in Cluster 2. Hence, we classified Bulgaria as belonging to Cluster 2.

Before proceeding with estimation of the model consisting of the clustered data, we test for the poolability of the respective clusters. Country dummies are constructed to aid poolability of the clustered data after which we test the hypothesis that the model has a homogeneous slope and therefore poolable. The hypothesis is constructed by comparing a panel "within" fixed effect model to a panel variable coefficient model. For both models, we specify a fiscal reaction function where the primary balance is a function of the lagged debt-to-GDP ratio and other control variables. Table 3.5 shows that for Cluster 1, we cannot reject the null hypothesis of model stability considering the p-value. Similarly, the null of model stability cannot be rejected for Cluster 2, implying that both the pooled coefficient model and variable coefficient model are comparable. The model for both datasets/clusters is

[3] We estimated the optimal number of clusters which turned out to be 4. However, using 4 clusters implies less number of observations for each group. Given that penalized splines require quite a number of observations for efficient estimation, we use 2 clusters in order to avoid loss in degrees of freedom.

Table 3.5: Panel poolability test - clusters

	Cluster 1				***Cluster 2***			
Variables	F-stats.	P-value	Df1	Df2	F-stats	P-value	Df1	Df2
Values	0.837	0.715	35	72	0.943	0.650	161	190
Num. Obs	120				380			

Chow test for poolability of panel data. Fixed effects "within model" and pooled panel model are compared to panel variable coefficient model. The null hypothesis states model is stable, and hence both models are comparable. The alternative hypothesis implies model instability. The estimate model is as follows $S_{i,t} = \mu_i + \alpha X_{i,t-1} + \sum_{j=1}^{m} \beta_j Y_{it}^j + \sum_{k=1}^{s} \theta_k D_{it} + v_{i,t}$, where D_{it} represent a vector of country dummies for cluster 1 consisting of Greece, Italy and Portugal. Country dummies constructed for cluster 2 are Ireland, Germany, Czechia, Finland and Spain.

poolable and hence, estimation and inference can be reliably made since both cluster datasets are poolable.

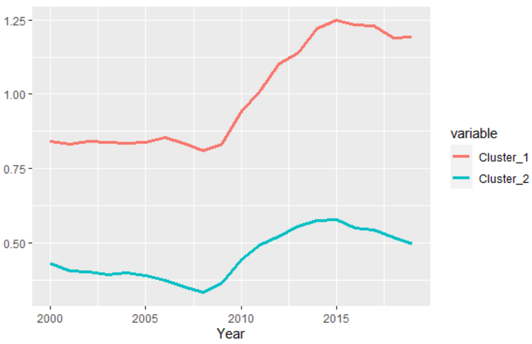


Figure 3.1: Debt ratio - clusters

Figure 3.2: Primary balance ratio - clusters

Figure 3.1 and 3.2 depicts comparability (visuals) of the average debt-to-GDP ratio and the primary balance for the respective clusters. Cluster 1 has a relatively high average debt-to-GDP ratio as compared to Cluster 2. The average primary balance is comparable between the two clusters (figure 3.2), especially before 2010; however, significant difference can be seen especially after the 2010 European debt crisis, as Cluster 1 generated a higher primary balance comparably. It can be observed that the difference in clusters is driven mostly by the variability of the debt-to-GDP ratio.

We approach the estimation strategy in two ways; firstly, we proceed by estimating a linear model using a panel linear fixed effects estimator in order to obtain point estimates for our inference. Secondly, we go beyond linearity by applying panel splines to estimate the nonlinear effect of the fiscal reaction function. This allows us to visualize the behaviour of the reaction coefficient over the distribution of the covariates.

The upper half of table 3.6 depicts the linear reaction function using estimates from fixed effects models based on the estimation of equation (3.12). As mentioned in the preceding subsection, we construct country dummies for Greece, Italy and Portugal to aid the poolability of our parameters. Standard errors are estimated using so-called whitening of the residuals (Heteroscedasticity and Autocorrelation Consistent (HAC) covariance matrix estimation) according to Newey and West (1987) so that they are robust against serial correlation and heteroskedasticity. Hence, the true standard errors are reported in the table. Regarding the presence of cross-sectional dependence (CSD), we resort to the Pesaran test (see Pesaran (2004)) to test the residuals of the linear model for CSD. The null hypothesis of the test indicates the absence of CSD. Based on the reported p-value, we cannot reject the null at 5% significance level. However, rejection at 10% significant level is possible and hence, the evidence of CSD is not so strong in the model. The origin of cross-sectional dependence could be due to the fact that the countries in the panel are linked geographically and economically (EU), and hence, they respond to shocks in a similar way. It could also be due to spillover effects from shocks or policies from certain countries, for instance, which could have implications for other countries. This has the potential of distorting the standard errors (inconsistencies) due to residual correlation. Driscoll and Kraay (1998) proposed a non-parametric robust covariance matrix estimator that yields standard errors that are robust to general forms of cross-sectional dependence (also robust against heteroscedasticity). The fourth column in table 3.6 depicts the standard errors from the Driscoll and Kraay estimator (from henceforth known as SCC). We notice that estimates of the SCC standard errors are not substantially different from the standard errors according to the HAC variance-covariance estimator by Newey and West.

Regarding the regression coefficients, we observe a positive and statistically significant lagged debt coefficient indicating fiscal sustainability for "Cluster 1" (European countries with high debt ratio), implying the evidence that active fiscal policy reacts positively to increases in the debt-to-GDP ratio. The intuition is that if debt-to-GDP ratio increases, higher primary balances are generated to offset or pay for the increases in debt. A negative and statistically significant output gap coefficient implies evidence of pro-cyclical fiscal policy, which is to say, during recessions (negative output gap), a primary surplus is generated, whilst deficits are generated during booms or good economic times (where potential GDP exceeds actual GDP). Inflation has a positive and statistically significant impact on the primary surplus. The coefficient for the trade variable is also positive and statistically significant (if we account for CSD), indicating that trade openness increases the primary surplus. Finally, the country dummies do not connote much significant economic interpretation as they only aided with pooling of the econometric model.

As a robustness check, a model that considers both individual and time effect is estimated and presented in Table 3.11 in the appendix. Comparing results to the upper half of the main specification in Table 3.6 indicates similarities and that, the main story does not change significantly. Positive and statistically significant debt-to-GDP ratio coefficient provides evidence of sustainability of fiscal stance in cluster 1, which is confirmed by the model with "twoways effect".

The lower half of table 3.6 depicts the estimation of the nonlinear part of the model using a panel penalized spline estimator. To deal with potential serial correlation, we used the Feasible Least Square procedure (FLS) based on Cochrane-Orcutt approach. The idea is to estimate an autocorrelation coefficient based on the residuals and pre-multiply the resulting coefficient by the design matrices and the dependent variable in order to transform the residuals. This leads to a model with homoskedastic and uncorrelated errors (Puetz and Kneib, 2018). We find evidence of AR1 (Autocorrelation of order 1) process in the residuals (see figure 3.14 in the appendix). Hence we correct via FGLS approach discussed above. The estimated degree of freedom for lagged debt and output gap are both greater than one and statistically significant indicating a high degree of nonlinearity.

Table 3.6: Estimation - cluster 1

Panel Linear Model

Variable	*Estimate*	*S.E (HAC)*	*S.E (SCC)*
lagged debt	0.071	0.020***	0.014***
Output gap	-0.031	0.016*	0.009***
Inflation	0.501	0.267*	0.246**
Trade	0.180	0.120	0.093*
Dummy (Greece)	0.067	0.030**	0.035*
Dummy (Italy)	-0.032	0.032	0.021
Dummy (Portugal)	-0.035	0.024	0.020*
R-squared	0.595		
CSD test	2.07(0.038)		
Numb. Obs	120		

Panel nonlinear model (spline)

Variable	*edf*	*F-stat*
lagged debt	3.031***	16.024
Output gap	3.085**	2.832
Inflation	1.001	2.273
Trade	2.381	1.581
R-squared	0.637	
Numb. Obs	114	

To inspect the presence of potential CSD, the panel spline residuals are displayed in figure 3.10. The individual country residuals are quite unique and do not exhibit strong a correlation, and hence, not enough evidence of CSD. Next, we visualize the behavior of the reaction function. Figure 3.3 depicts the plot of the fitted penalized spline over the distribution of the lagged debt. We can observe an upward-sloping curve over the distribution of debt. The initial reaction of the primary balance as debt increases is fairly weak; however, the reaction becomes stronger with further increases in debt. The result supports the argument that higher debt warrants higher primary balance in order to ensure fiscal sustainability. Hence, there is evidence that this group of countries allow debt to grow before taking drastic measures in terms of adjusting their primary balance, for instance.

Table 3.7 presents estimates of results for Cluster 2 (countries with relatively low debt-to-GDP ratio). The upper half of the table depicts estimates of the panel linear model, whilst the lower part shows the non-parametric estimates. Once again, there is evidence of debt sustainability due to the positive and statistically significant lagged debt coefficient, whether we consider HAC or SCC standard errors. However, the reaction coefficient of the lagged debt variable (0.065) is relatively lower than that of Cluster 1 (0.071). The negative coefficient of the output gap also depicts the pro-cyclical nature of fiscal policy, but it is however, statistically insignificant. Inflation and trade both exert positive pressure on the primary balance. Once again, we present estimates of a model with both individual and time effects and show in the appendix (Table 3.12) that the result does not differ substantially, compared to the main specification.

Looking closely at the nonlinear section, lagged debt, inflation and trade variables are all statistically significant. Due to the evidence of AR1 process (see figure 3.15 in the appendix, which depicts the partial autocorrelation plots), we apply FGLS using the Cochrane-Orcutt technique as a second-step procedure to correct the autocorrelation in the residuals. Figure 3.4 presents the visualization of the fiscal reaction function for Cluster 2. It can be observed that the primary balance rises initially as debt increases, fluctuates and remains fairly flat at a very high level of debt. The residuals of the panel splines also do not exhibit strong evidence of CSD as can be seen in figure 3.11, since most of the individual country residual plots appear unique.

Obviously, the response function is different in the two clusters: while the response in the case of Cluster 1 is almost monotonously increasing with higher public debt-to-GDP ratios, Cluster 2 with the smaller economies resembles a partial inverse U-shaped form. This declining and fairly flat reaction of the primary balance for high values of the debt ratio is similar to the 'fiscal fatigue' argument as suggested by Gosh et al. (2013). For both cluster specifications, the results in table 3.6 and table 3.7 mainly support the general findings that active fiscal policy reacts in a positive manner to increases in debt, indicating evidence of sustainable fiscal policy. There is also evidence of pro-cyclicality of fiscal policy for both clusters. Inflation and trade were also found to affect the primary balance positively.

3.4.5 Crisis and non-crisis period

In what follows, we split the sample in order to study the behavior of the fiscal reaction function in good economic period and crisis periods. Regarding the crisis period, we choose the period between 2008 and 2012 which coincides with the global financial crisis and European debt crisis. Further, in those years, most of the EU countries were under the excessive deficit procedure due to violation of the Maas-

Table 3.7: Estimation - cluster 2

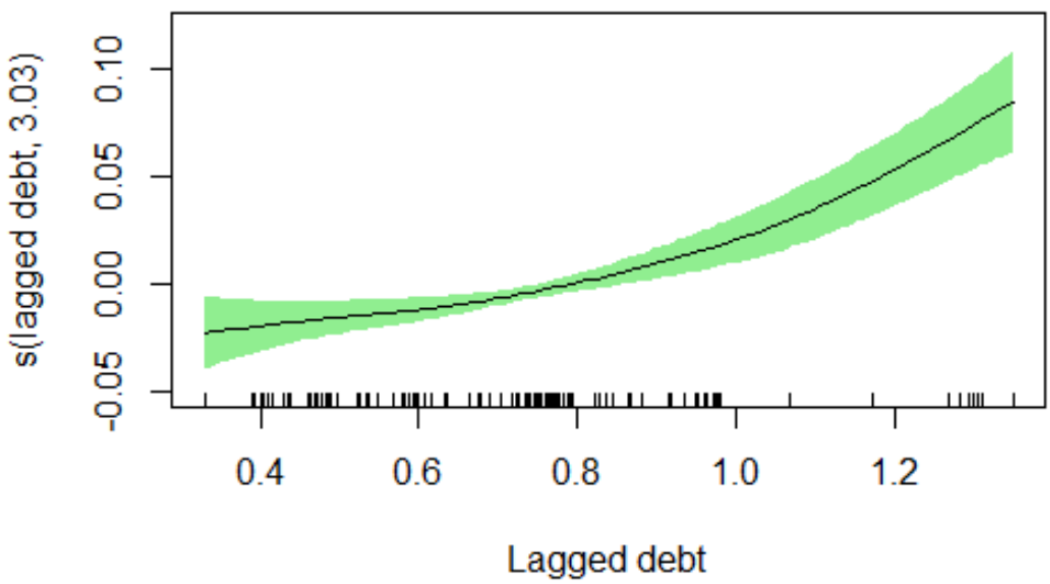
Panel Linear Model

Variable	*Estimate*	*S.E (HAC)*	*S.E(SCC)*
lagged debt	0.065	0.011***	0.012***
Output gap	-0.006	0.007	0.007
Inflation	0.154	0.056***	0.065**
Trade	0.123	0.032***	0.030***
Dummy (Ireland)	-0.107	0.029***	0.023***
Dummy (Germany)	-0.015	0.032	0.033
Dummy (Czechia)	0.115	0.040***	0.041***
Dummy (Finland)	-0.213	0.047***	0.042***
Dummy (Spain)	-0.052	0.029*	0.020**
R-squared	0.19		
CSD test	6.105(0.000)		
Numb. Obs	380		

Panel nonlinear model (spline)

Variable	*edf*	*F-stat*
lagged debt	4.686***	4.066
Output gap	1.001	1.476
Inflation	4.373***	7.530
Trade	6.570***	8.445
R-squared	0.293	
Numb. Obs	361	

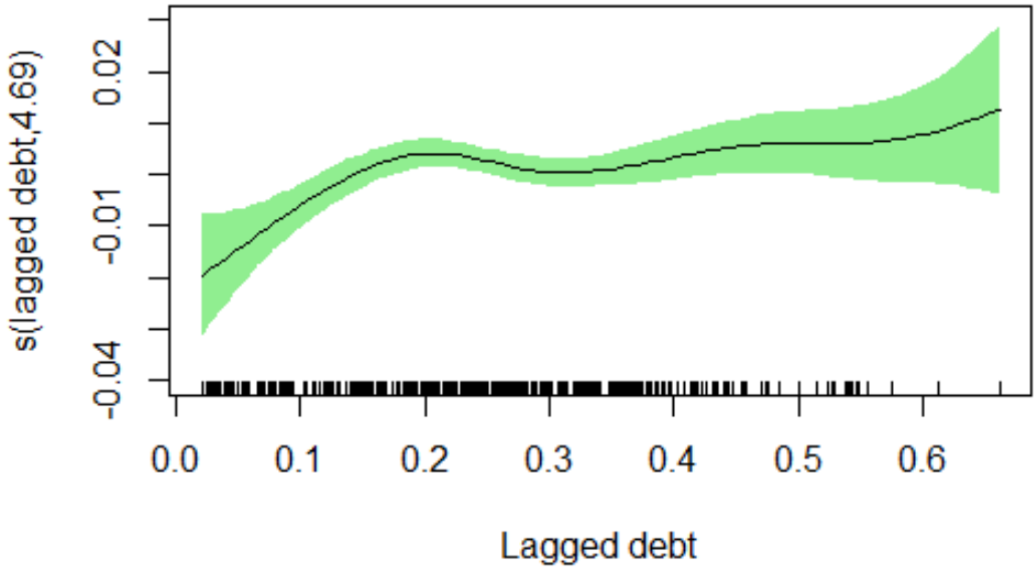
Figure 3.3: Cluster 1



tricht Treaty criteria.[4] Due to the low number of observations for Cluster 1, we are unable to split the data further, as we have minimal degree of freedom. Hence, clus-

[4]See European Commission (2021) for information regarding the excessive deficit procedure.

Figure 3.4: Cluster 2



tering into crisis and non-crisis is based on Cluster 2 data (subsample of countries with relatively low debt). The crisis period span a period of 5 year. The non-crisis period spans from 2000 to 2007 and then from 2013 to 2019. This period is characterised by relatively good economic times without major macroeconomic shocks. Once again, to aid poolability, we construct country dummies to augment the model for only non-crisis period since crisis sample is poolable without country dummies. Table 3.8 presents the results of the pooling test for both crisis and non-crisis periods. It can be seen that based on the p-values of the chow test, the coefficient or slope parameters are poolable in both cases of crisis and non-crisis periods.

Table 3.8: Panel poolability test : crisis and non-crisis

	Crisis				***Non-crisis***			
Variables	*F-stas.*	*P-value*	*Df1*	*Df2*	*F-stats*	*P-value*	*Df1*	*Df2*
Values	0.638	0.901	54	19	0.641	0.994	162	95
Num Obs	120				380			

Chow test for poolability of panel data. Fixed effect "within model" and pooled panel model are compared to panel variable coefficient model. The null hypothesis states model is stable and hence both models are comparable. The alternative hypothesis implies model unstability. The estimated model is as follows $S_{i,t} = \mu_i + \alpha X_{i,t-1} + \sum_{j=1}^{m} \beta_j Y_{it}^{j} + \sum_{k=1}^{s} \theta_k D_{it} + v_{i,t}$, where D_{it} represent a vector of country dummies for non-crisis period comprising of Spain, Cyprus, Ireland, Estonia and Bulgaria.

The estimates for the reaction function for crisis and non-crisis periods can be found in table 3.9 and table 3.10 respectively. Regarding the crisis period, the reaction coefficient is positive, but statistically insignificant irrespective of whether we consider HAC standard errors or SCC (correcting for cross-sectional dependence) with reference to table 3.9. Moreover, with the exception of trade (if we consider SCC standard errors), the remaining control variables are statistically insignificant, which is not very surprising given that the period under consideration corresponds to an economic crisis period. There is evidence that trade drives primary surplus even during crisis period. Generally, we infer that during economic crisis period, there is lack of evidence of sustainable fiscal behaviour according to our sample of study due to statistically insignificant results. The lower panel of table 3.9 depicting the non-parametric estimations (panel splines) reveals that only lagged debt and inflation are statistically significant. The pacf plot in the appendix (figure 3.16) indicates the presence of AR1 process. Hence, we correct using FGLS approach as described in the previous subsection.

Table 3.9: Estimation - crisis period

	Panel Linear Model		
Variable	*Estimate*	*S.E (HAC)*	*S.E(SCC)*
lagged debt	0.061	0.049	0.051
Output gap	0.009	0.007	0.006
Inflation	0.063	0.089	0.090
Trade	0.074	0.122	0.040*
R-squared	0.067		
CSD test	2.439(0.015)		
Numb. Obs	95		

	Panel nonlinear	model (spline)	
Variable	*edf*	*F-stat*	
lagged debt	3.228***	5.681	
Output gap	1.850	0.865	
Inflation	3.179*	2.628	
Trade	4.454	5.460	
R-squared	0.451		
Numb. Obs	76		

The estimates from table 3.10 (non-crisis period) present a slightly different picture as compared to crisis sample. The debt reaction coefficient is positive and statistically significant, implying evidence of a sustainable fiscal stance during non-crisis period. The negative (and significant) output gap indicates pro-cyclicality. The inflation coefficient is statistically insignificant, whilst the trade variable is positive and significant, implying that trade drives fiscal surplus. Once again, the CSD test indicates the presence of residual CSD. Hence, we present a report of standard errors (SCC) which are robust against CSD. The lower panel of table 3.10 presents the non-parametric estimates of the model. We once again resorted to FGLS approach using Cochrane-Orcutt procedure due to the presence of serial correlation in the residuals (See results of the pacf in figure 3.17 in the appendix). Lagged debt, Inflation and trade variables are statistically significant and, moreover have a degree of curvature based on the edf and demonstrate that, the relationship between the primary surplus ratio, and the explanatory variables is characterized by nonlinearities.

Figure 3.5 depicts the shape of the smooth function for the crisis specification. It can be noticed that the reaction coefficient increases initially, flattens at fairly high debt levels and then trends upwards at very high debt levels which is synonymous with the behaviour of the reaction coefficient for Cluster 1. Figure 3.6 however, depicts a different story. For non-crisis periods, the reaction coefficient exhibits fiscal fatigue behaviour. An initial strong response of the primary balance to increases in debt level. However, as debt reaches a particular threshold, the response becomes weak, wiggly and eventually falls, which can be likened to an inverted U-shape. During normal times (or good economic times) and without external support or austerity measures, allowing debt to grow beyond a certain threshold could pose a fiscal sustainability risk as we have seen in Figure 3.6. It becomes difficult to raise the primary balance at very high debt levels, hence an indication of the fiscal fatigue narrative. Finally, the individual country residual plots (panel splines) shown in figure 3.12 and figure 3.13 for non-crisis and crisis samples do not indicate strong evidence of CSD since the residuals appear unique.

Table 3.10: Estimation : non-crisis period

	Panel Linear Model		
Variable	*Estimate*	*S.E (HAC)*	*S.E(SCC)*
lagged debt	0.046	0.011***	0.010***
Output gap	-0.010	0.006*	0.005**
Inflation	0.015	0.035	0.025
Trade	0.0001	0.0000***	0.0000***
Dummy (Spain)	-0.083	0.017***	0.018***
Dummy (Cyprus)	0.171	0.046***	0.046***
Dummy (Ireland)	-0.084	0.013***	0.013***
Dummy (Estonia)	-0.309	0.038***	0.030***
Dummy (Bulgaria)	0.131	0.050**	0.063**
R-squared	0.308		
CSD test	2.207(0.027)		
Numb. Obs	285		

	Panel nonlinear	model (spline)
Variable	*edf*	*F-stat*
lagged debt	5.610***	5.586
Output gap	1.005	1.814
Inflation	1.002*	3.108
Trade	6.052***	9.676
R-squared	0.372	
Numb. Obs	266	

Summing up our additional results, some interesting findings exist beyond the linear fixed effects estimation. By distinguishing two (data-driven) clusters, it turns out that the fiscal response differs for each cluster. For the cluster with mostly

Figure 3.5: Crisis

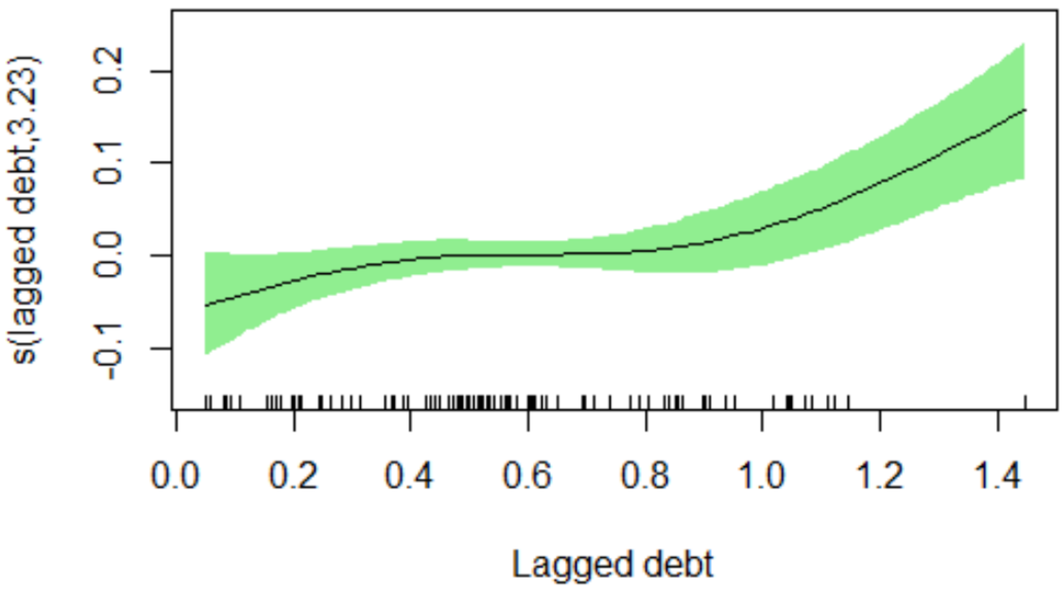
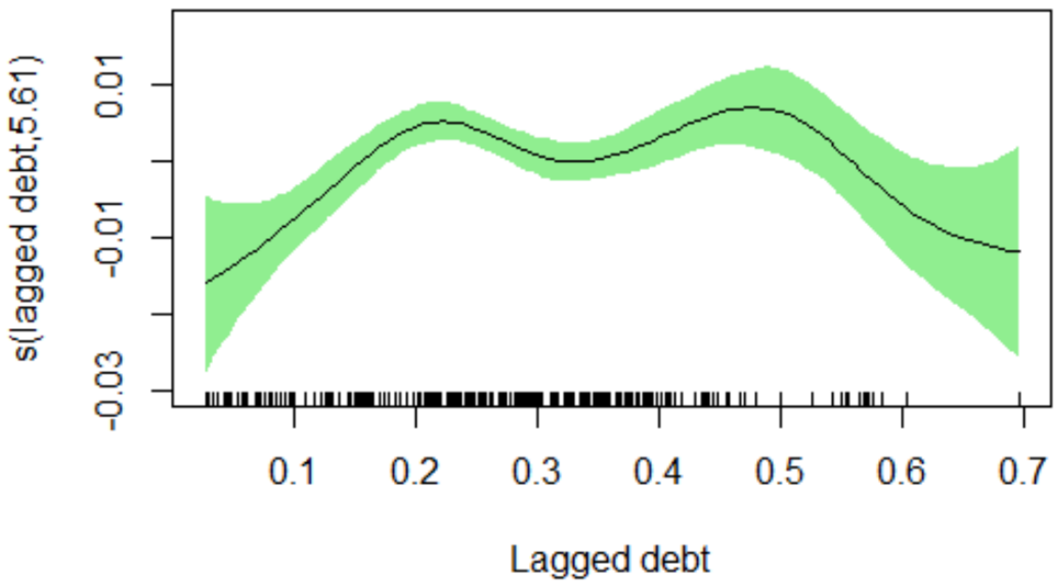


Figure 3.6: Non-crisis



small countries, in particular, 'fiscal fatigue' characteristics appear, meaning that the response to higher debt peters out and, finally, decreases or remains flat as the debt-to-GDP ratios become larger. For the cluster containing higher debt-to-GDP ratios, the response to higher debt is a monotonically increasing function of the debt-to-GDP ratio. Moreover, in non-crisis years, there is even stronger evidence of fiscal fatigue, which is to say that, higher debt beyond a certain threshold induces a burden on fiscal policy, making it challenging to raise the primary balance in response to increases in debt. However, during crisis years, the response (behavior of the reaction function) is monotonously increasing, meaning a stronger reaction to higher debt if governments face large debt ratios. This can be justified by the implementation of austerity measures during crisis period to avoid default. However, the point estimates for the crisis period is statistically insignificant.

Comparing these results with previous studies indicates some similarities despite the different approaches. For instance, Gosh et al. (2013), with a different time period (1970-2007) and different sample (23 advanced economies, including countries outside Europe) and different empirical framework, found evidence of fiscal fatigue with a turning-point of debt ratio level of about 90-100%. Our estimations for Cluster 2 (countries with low debt ratios) and non-crisis sub-samples yields similar results of fiscal fatigue. Fournier and Fall (2015) also find fiscal fatigue behaviour (OECD economies, 1985:2013, threshold model) for a debt ratio of about 170% of GDP and for the Euro area group (15 countries) with relatively lower turning points (152% and 167%). On the other hand, Checherita-Westphal and Zdarek (2017) find only weak support for fiscal fatigue (18 Euro area) resorting to an approach which takes actual fiscal behaviour into account.

3.5 Conclusion

This paper studied debt sustainability analysis by resorting to panel linear fixed effects and penalized panel spline technique. Based on data for 25 EU economies from 2000 to 2019, we estimated the fiscal response function by analyzing the relationship between the discretionary fiscal policy in terms of the cyclically adjusted primary balance-to-GDP ratio and the lagged debt-to-GDP ratio. A positive coefficient on average, indicates a sustainable public debt policy, which is supported by our results. However, we have seen that the relationship is not homogeneous across the distribution of the debt ratios but, rather, varies with the size of the debt ratio when we employ a panel penalized spline estimator.

We used a data-driven algorithm (k-means) to cluster the data set into two distinct groups. Results for the first cluster, consisting of economies with higher debt ratio, show a strong response of the primary balance to increases in the debt ratio yielding strong evidence of fiscal sustainability. Regarding the second cluster (countries with relatively low debt ratios), the average reaction coefficient takes on a relatively lower value and the smooth function is characterized by a partially inverted U-shaped relationship: it rises with increasing debt-to-GDP ratios, reaches a maximum and then, remains fairly flat. Additionally, both clusters are characterized by pro-cyclicality of fiscal policy due to negative output gap coefficient.

Furthermore, the dataset is classified into two categories to ascertain the reaction function and behaviour of fiscal policy in times of financial/debt crisis and in normal times. The reaction function in normal times exhibits a similar pattern as compared to Cluster 2 with stronger evidence of fiscal fatigue, meaning that the effort of counter-steering peters out and declines once the debt-to-GDP ratios exceed a certain threshold. This supports the argument that allowing debt to grow beyond a certain threshold induces some kind of fiscal pressure, making it difficult to raise primary balance in response to growing debt. While during years of crisis, the reaction of the primary balance to public debt and some control variables have been found to be statistically insignificant, pointing to the effects of adverse economic period on the macro-economy.

These refinements are particularly important for policy implications and help improve fiscal recommendations as they indicate that the current debt status and the economic situation is essential for the assessment and evaluation of sustainability. Our results show that, yes, the size of the debt does matter for the fiscal response. The level of the explanatory variable influences the reaction coefficient and, thus, fiscal sustainability. Furthermore, the fiscal reaction function in crisis period differs significantly from non-crisis period. Policy recommendations need to consider the status of the current debt situation in order to be successful, as the reaction function shows different behavior for low debt levels compared to medium and high debt ratios. Besides, it differs based on the macroeconomic conditions. Our findings align with other literature contributions that focus on nonlinear fiscal behaviour and the study of "fiscal fatigue" in debt sustainability analysis.

3.6 Appendix

Table 3.11: Estimation - cluster 1 with both individual and time effects

Panel Linear Model			
Variable	*Estimate*	*S.E (HAC)*	*S.E (SCC)*
lagged debt	0.100	0.022***	0.022***
Output gap	-0.040	0.017**	0.010***
Inflation	-0.058	0.221	0.169
Trade	-0.014	0.092	0.089
Dummy (Greece)	0.061	0.024**	0.030**
Dummy (Italy)	-0.005	0.039	0.027
Dummy (Portugal)	-0.016	0.021	0.025
R-squared	0.618		
CSD test	-3.125(0.002)		
Numb. Obs	120		

Table 3.12: Estimation - cluster 2 with both individual and time effects

Panel Linear Model			
Variable	*Estimate*	*S.E (HAC)*	*S.E (SCC)*
lagged debt	0.063	0.010***	0.013***
Output gap	-0.004	0.008	0.007
Inflation	0.073	0.041*	0.048
Trade	0.184	0.036***	0.035***
Dummy (Ireland)	-0.098	0.025***	0.022***
Dummy (Germany)	0.048	0.040	0.036
Dummy (Czechia)	0.159	0.040***	0.037***
Dummy (Finland)	-0.180	0.028***	0.029***
Dummy (Spain)	-0.050	0.022**	0.017***
R-squared	0.212		
CSD test	-2.106(0.035)		
Numb. Obs	380		

Table 3.13: Summary statistics - cluster 1

Variable	Mean	Std dev	Min	Max	25th perc	75th perc
Debtratio	0.993	0.317	0.456	1.862	0.726	1.105
Primary balance	0.006	0.0318	-0.097	0.097	-0.009	0.021
Output gap	0.000	0.048	-0.460	0.689	-0.045	0.037
Inflation	0.017	0.013	-0.020	0.054	0.010	0.023
Trade	-0.010	0.147	-0.127	.056	-0.032	0.028

Figure 3.7: K-mean plot (Standardized variables)

Table 3.14: Summary statistics - cluster 2

Variable	Mean	Std dev	Min	Max	25th perc	75th perc
Debtratio	0.456	0.234	0.038	1.199	0.280	0.613
Primary balance	-0.002	0.030	-0.277	0.0783	-0.015	0.016
Output gap	0.000	0.167	-1.066	0.962	-0.034	0.028
Inflation	0.030	0.041	-0.097	0.432	0.011	0.038
Trade	0.033	0.094	-0.206	.360	-0.016	0.067

Table 3.15: Summary statistics - crisis

Variable	Mean	Std dev	Min	Max	25th perc	75th perc
Debtratio	0.432	0.232	0.038	1.110	0.273	0.617
Primary balance	-0.014	0.041	-0.277	30.043	-0.033	0.043
Output gap	0.008	0.216	-1.066	0.962	-0.020	0.039
Inflation	0.0214	0.029	-0.097	0.160	0.009	0.033
Trade	0.032	0.089	-0.197	0.329	-0.014	0.057

Table 3.16: Summary statistics - non-crisis

Variable	Mean	Std dev	Min	Max	25th perc	75th perc
Debtratio	0.465	0.234	0.046	1.199	0.282	0.608
Primary balance	0.002	0.024	-0.087	0.078	-0.011	0.017
Output gap	-0.002	0.147	-0.702	0.891	-0.037	0.024
Inflation	0.034	0.045	-0.006	0.432	0.013	0.039
Trade	0.034	0.096	-0.207	0.360	-0.017	0.068

Table 3.17: Correlation matrix

Variable	Trade	Inflation	Output gap	Lagged debt	Primary balance
Trade	1				
Inflation	-0.27	1			
Output gap	-0.07	0.07	1		
Laggedebt	-0.10	-0.31	-0.12	1	
Primary balance	0.26	-0.08	-0.1	0.21	1

Figure 3.8: Debt-to-GDP ratio (Individual countries)

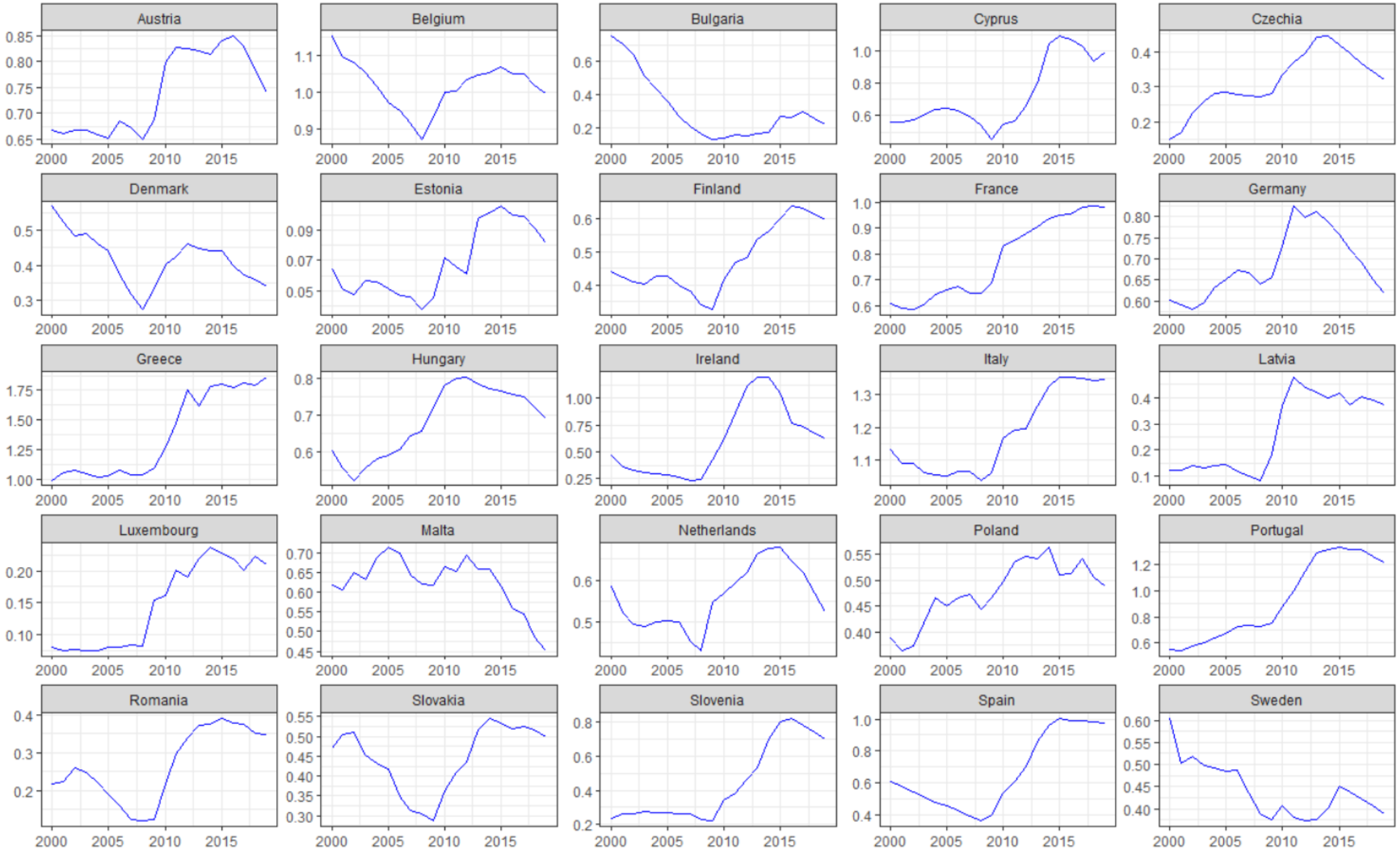


Figure 3.9: Primary balance-to-GDP ratio (percentage)

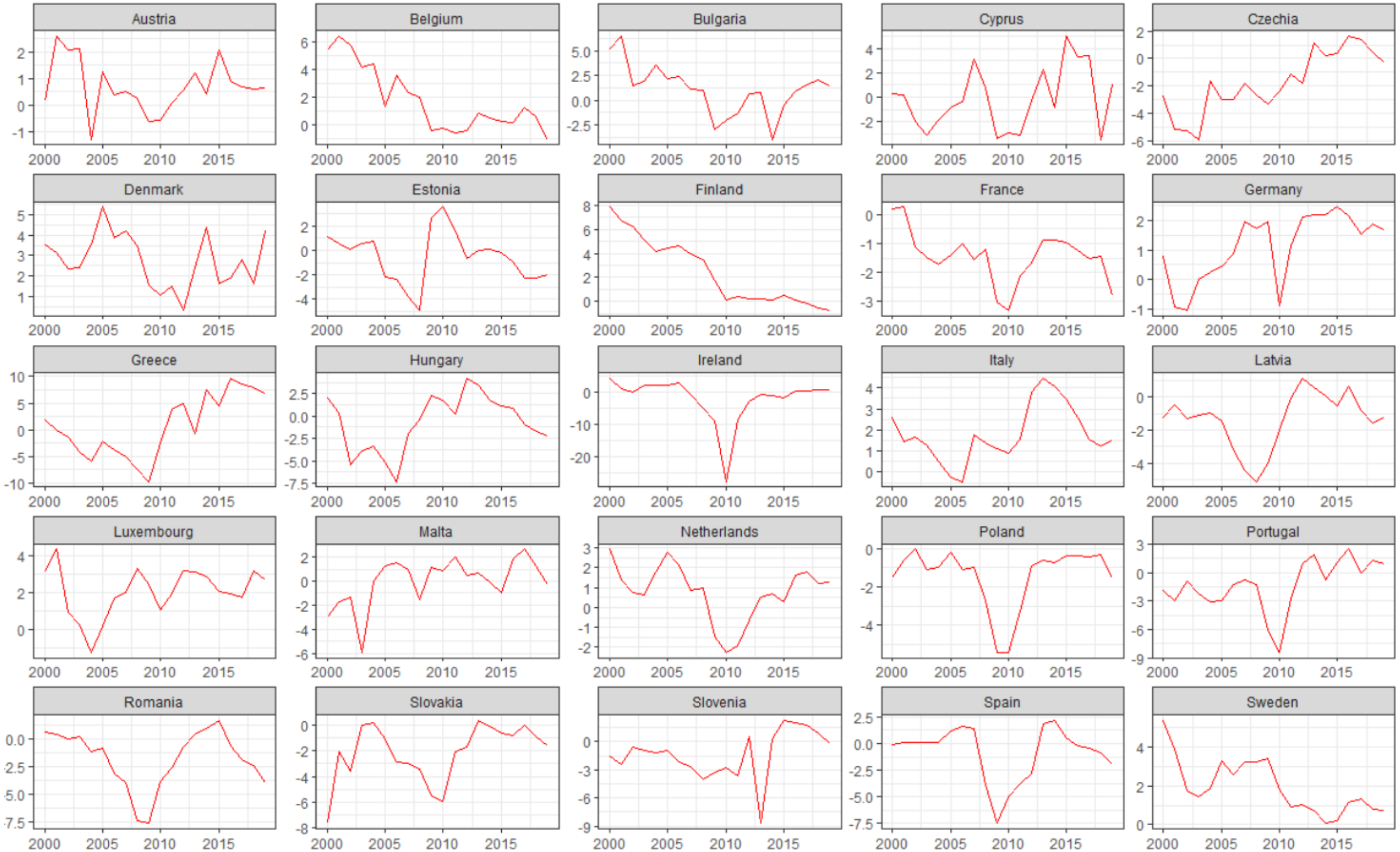


Figure 3.10: Residual plot - cluster 1 (spline)

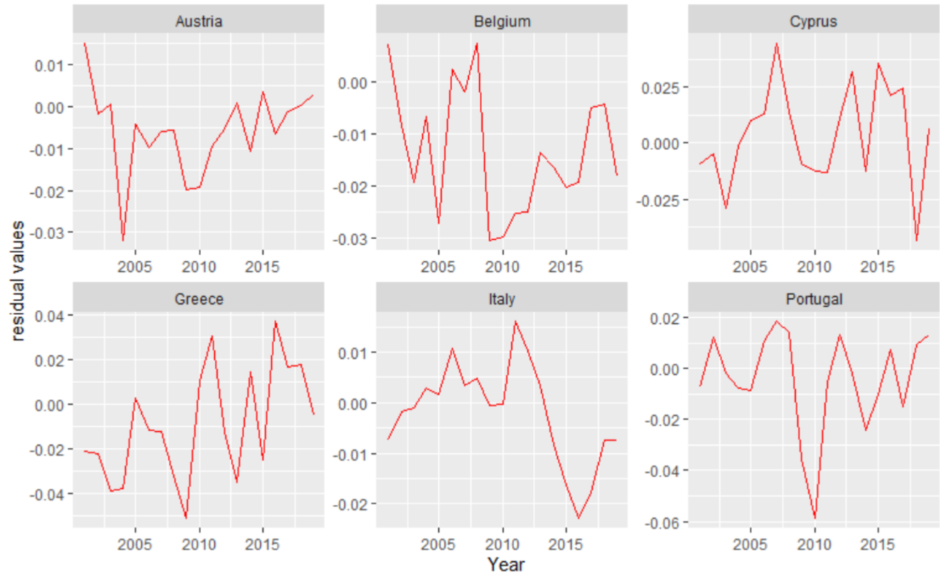


Figure 3.11: Residual plot - cluster 2 (spline)

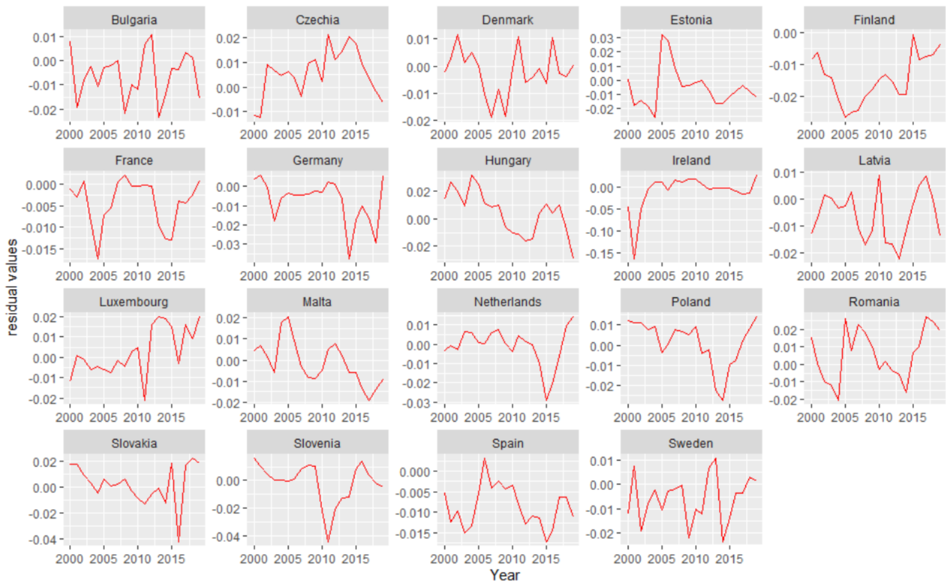


Figure 3.12: Residual plot, non-crisis period (spline)

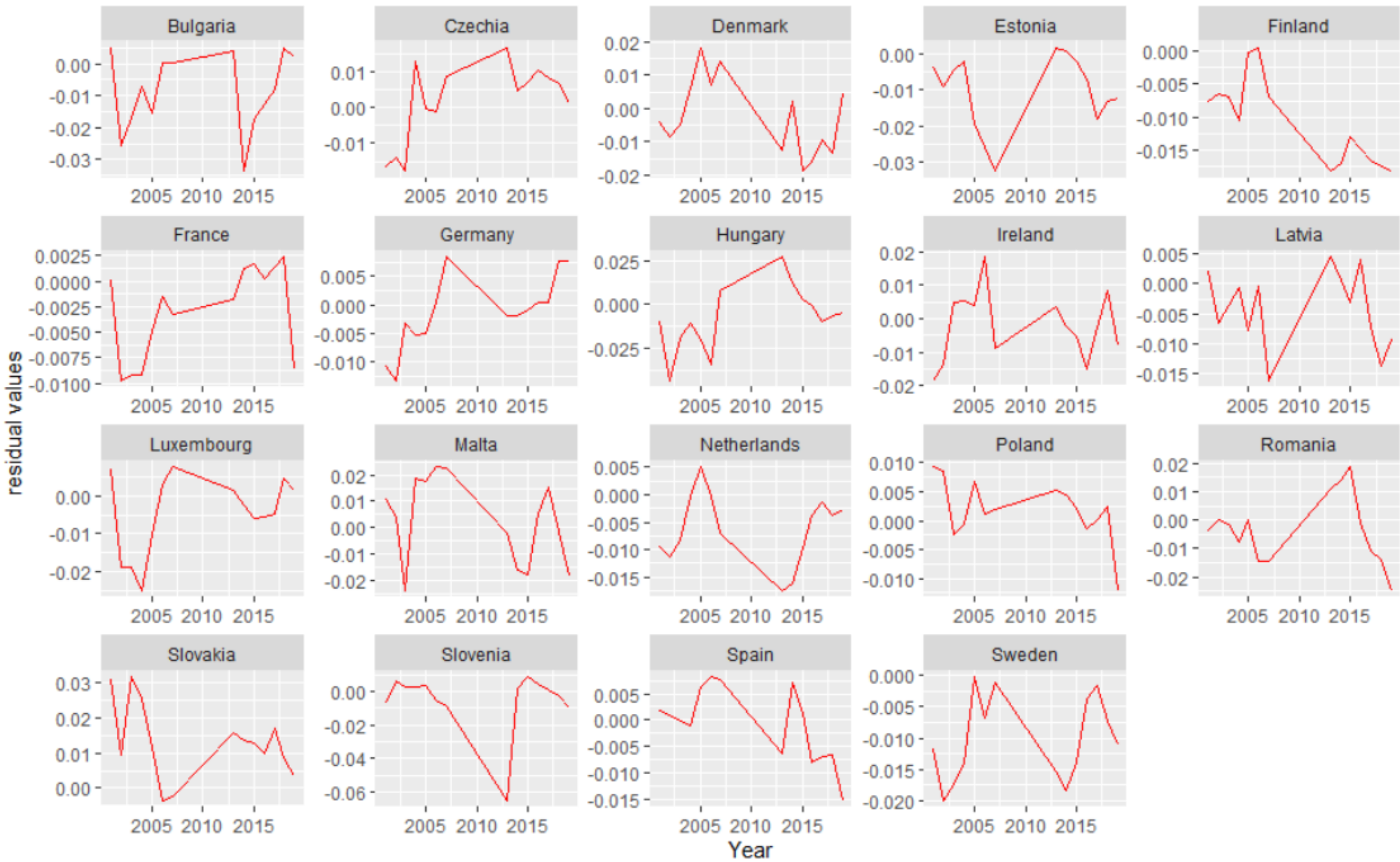
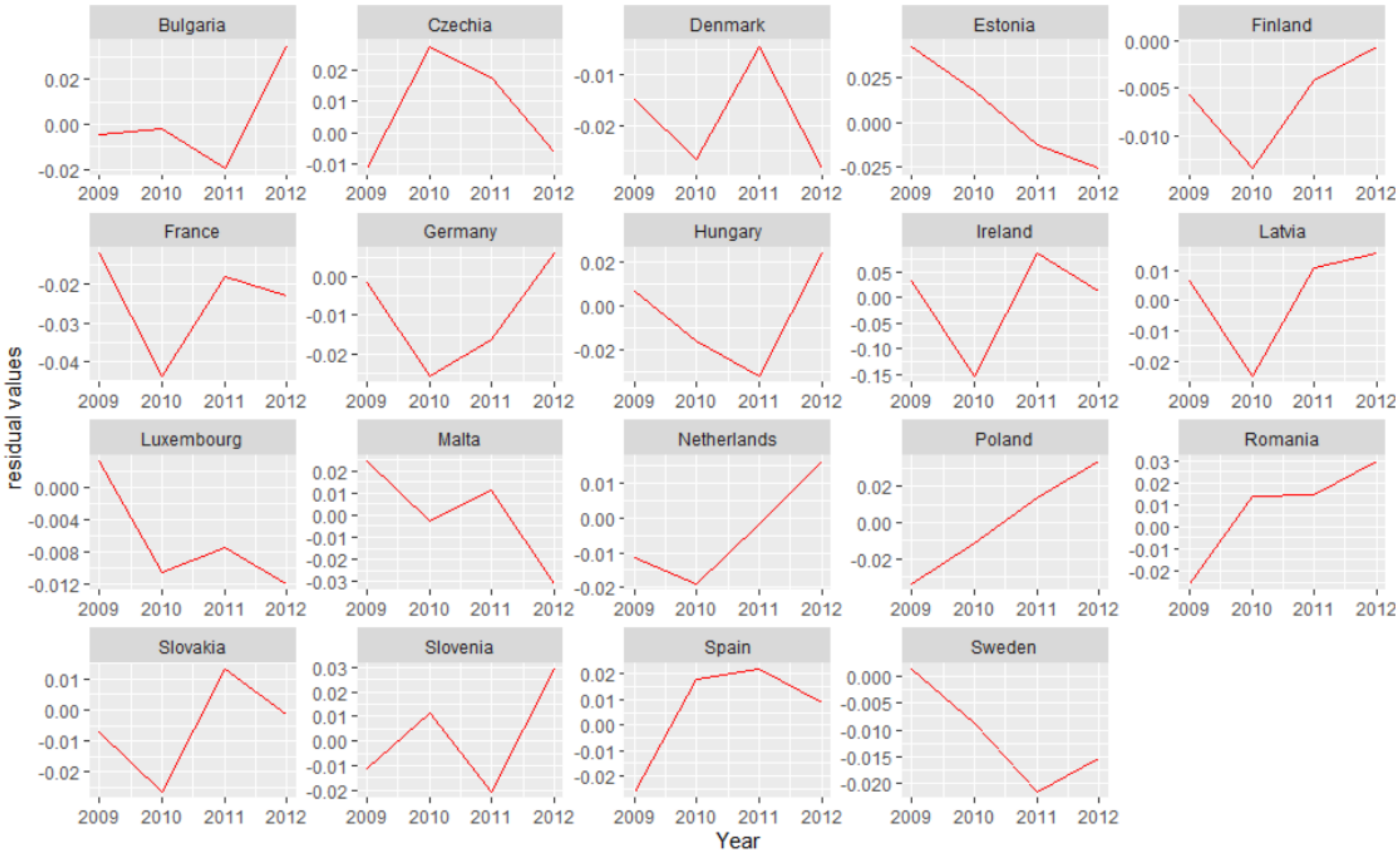


Figure 3.13: Residual plot - crisis period (spline)



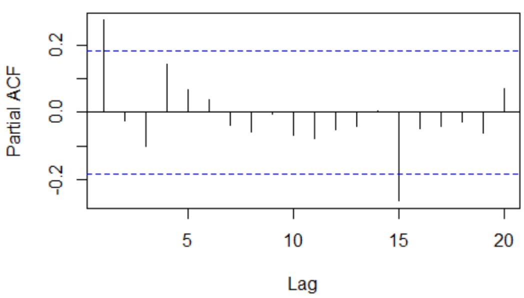


Figure 3.14: Partial ACF - cluster 1

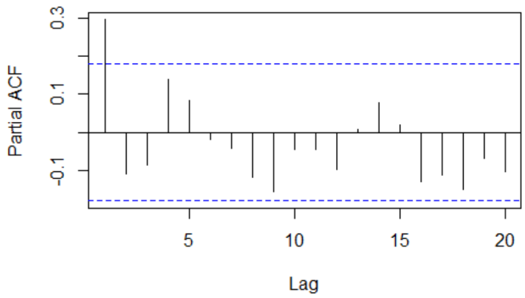


Figure 3.15: Partial ACF - cluster 2

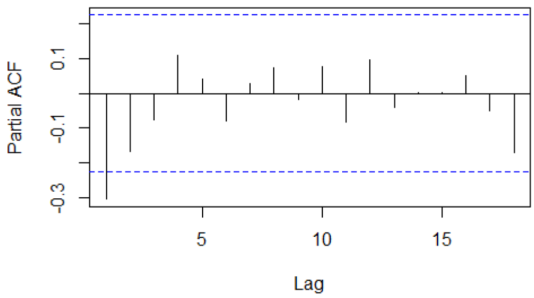


Figure 3.16: Partial ACF - crisis

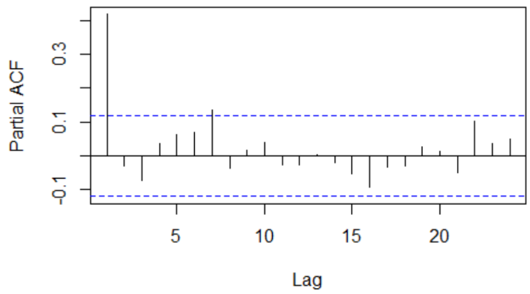


Figure 3.17: Partial ACF - non-crisis

3.6.1 Data availability and scripts

Data used for this paper is available on my Github page (https://github.com/Benjamin-Owusu/Thesis). The data was processed and analyzed with R statistical software. Accessibility to the data and scripts will be provided by the author upon reasonable request.

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Chapter 4

Regime-based debt sustainability analysis: evidence from euro area economies

4.1 Introduction

Debt sustainability discussions revived with the European debt crisis starting in 2008/2010 and have received additional attention due to the recent Covid-19 pandemic forcing governments worldwide to stabilise their economies with immense recovery programs primarily financed by credits, amid situations already characterised by high-debt ratios. Thus, analyzing the sustainability of fiscal policies and the budget positions is an essential task. Government interventions by way of fiscal policy have implications for the present budget and future decisions. Fiscal policy is one of the most powerful instruments in the design of public policy to achieve macroeconomic goals. Hence, to be able to conduct independent government spending and tax decisions, debt sustainability analysis is an important consideration.

Introduced by Bohn (1995, 1998), the fiscal response function, which estimates the reaction of the primary balance to changes in public debt relative to the gross domestic product (GDP), is a well-established approach to asses debt sustainability. Suppose the government reacts to a higher debt ratio by actively adjusting its discretionary fiscal policy in terms of a higher primary surplus. In that case, the fiscal policy under consideration is considered as sustainable. That relationship is commonly tested empirically in a single equation regression model and, more recently, in a standard panel set-up. The theoretical background and a formal model for debt sustainability analysis and the fiscal response function can be found in Greiner and Fincke (2015), for instance.

There have been several applications of this approach in the economics literature. For example, Afonso (2005), Afonso and Jalles (2019) and Beqiraj et al. (2018) presented a comprehensive overview of the literature. The fiscal response function has been applied in many ways according to the purpose of study and the particular research question. It may be related to public finances/ sustainability aspects (see Bohn (1995, 1998), Afonso (2005), Greiner and Fincke (2015) for example), cyclicality or business cycle studies (see Lane (2003), Kaminsky et al. (2004), Golinelli and Momigliano (2006), Gootjes and de Haan (2022) and Larch et al. (2021)), institutional aspects including government efficiency, rules and regulation (see Hallerberg and Wolff (2008), de Haan and Sturm (1994) for instance) and also political economy ideas (such as election effects etc., see Afonso (2008) for example). Golinelli and Momigliano (2006) provided a similar distinction regarding the four types of fiscal studies. The particular focus is then reflected in the set-up of the study, the collection of included variables, the emphasis of the research and the interpretation.

The aim of our study and the research question that we focus on belongs to the first type of literature contributions, namely on the influence of the level of debt on fiscal behaviour which is represented by the cyclically adjusted primary balance. However, most papers analyse linear single equation models or standard fixed effects panel regression models. With this paper, we research whether there are nonlinearities in the reaction of governments to rising public debt, implying that the level of the explanatory variable influences the relationship and the shape of the response. Particularly, we find that the reaction of the primary balance does vary with the size of the debt ratio, meaning that in situations with low-debt-to-GDP ratios the response turns out to be different from that in situations with high-debt ratios.

Some studies allowing for nonlinearities and heterogeneity in debt sustainability analysis can be found in the economics literature. For instance, the notion of "fiscal fatigue" has been introduced by Gosh et al. (2013), who detect a reversal in the behaviour of the primary balance as debt ratios become very high and the response peters out and becomes negative. Other applications in this direction are by Checherita-Westphal and Zdarek (2017), who found (weak) sustainability for eurozone members, as well as Fournier and Fall (2015), indicating high debt limits for OECD countries, Legrenzi and Milas (2012) find evidence of sustainability for highly indebted European economies and, Owusu et al. (2021) provide nonlinear empirical evidence of debt sustainability in the EU based on panel splines for instance. We contribute to this line of research and extend it by resorting to the estimation technique of panel smooth transition regression (PSTR) applied to euro area economies. PSTR has been established by Gonzalez et al. (2017) and allows to detect a threshold in the reaction function. It refrains from a country-wise perspective and applies a regime-switching model to detect nonlinearities in the data-generating process. Data is segregated into different regimes endogenously via a logistic regression function. Legrenzi and Milas (2012) applied similar approach based on single equation / country studies for a group of highly indebted economies, made up of Greece, Italy, Ireland, Portugal and Spain, based on historical time series, and found country-specific thresholds (varying with the debt level). Unlike Fournier and Fall (2015), we do not calculate country-specific debt limits but rather focus on the data-driven approach to determine different regimes (for fiscal behavior).

With the PSTR, we analyse the fiscal response that overcomes the pooling problem in the panel data context, where one coefficient fits all. The advantage is obvious: the number of regimes is determined by the data and the coefficients are estimated for each regime (for example, high or low-debt). In addition, they are not determined by particular economies, but by the respective debt situations. Individual country data points are not restricted to staying in the same group or category but can switch between the groups depending on the heterogeneity of their fiscal behaviour. Hence, we refrain from the country-wise classification of regimes. For instance, based on the value of the debt (transition variable), a country with a heterogeneous distribution of debt values could have data points in both high and low-debt regimes depending on the threshold value. This implies that different responses are feasible in the same country according to the different debt situations, i.e. the response is different in low-debt situations compared to high-debt situations. Therefore, the segregation of data into groups or regimes is not based on country-wise characteristics but rather based on individual data points.

Comparing PSTR to other similar panel regime-switching models, such as the Panel Threshold Regression (PTR) model developed by Hansen (1999), the PTR segregates the data points into groups/regimes with sharp thresholds or borders whilst the PSTR allows the regression coefficient to change smoothly from one regime to another based on a transition variable (Gonzalez et al., 2017). Specifically, in the PSTR setting regression coefficients are allowed to switch smoothly between two regimes characterised by low or high values of the transition variable. Smooth transitioning from one regime to another is feasible and can be justified by the fact that (except maybe in times of crisis), the effect of policy decisions is usually expected to have a gradual impact over a period of time rather than an abrupt impact. A case in point is the interest rates smoothing by most central banks. Therefore, PSTR models are more appealing and reflect reality better than other threshold or regime-based panel models. Previously, there has been limited application of regime-switching or nonlinear models to debt sustainability in the panel data context. There have been similar studies by Legrenzi and Milas (2012), however, they estimated multivariate time series data with the application of smooth transition regression and hence made inferences based on individual country data. Other papers that studied regime-switching models in the debt sustainability context include but are not limited to Ricci-Risquete et al. (2016), who analysed regime dependency of fiscal policy in Spain. Afonso and Toffano (2013) assessed the existence of fiscal policy regime shifts in the UK, Germany and Italy individually. Afonso et al. (2011) estimated the changes in fiscal policy regimes in Portugal. Using US and French data, Aldama and Creel (2016, 2019) resorted to regime-switching model-based sustainability tests for fiscal policy in each country. All the above papers focused on individual country studies and not on a group of countries. In contrast to that, we draw inferences for the entire euro area, which can be justified by the fact that in a currency union each government is committed to abiding by some well-defined rules (such as the Maastricht Treaty) that limit their fiscal behavior. Secondly, the previous authors applied Markov-switching models where the segregation into regimes (switching) is based on probabilities and, therefore, a stochastic process (Martin et al. (2013)). Contrarily, switching in the panel smooth transition regression framework is deterministic. It allows us to condition the model based on a known transition variable (such as the debt-to-GDP ratio, for instance) so that we are able to segregate the regimes based on a certain threshold value of the transition variable. Hence, with the PSTR estimation, we draw inferences based on the threshold value separating the regimes. The threshold value that separates the regimes has an economic meaning and therefore, the regimes are interpretable, for instance, a low-debt and a high-debt regime.

Our study contributes to the literature on the application of panel data for debt sustainability studies using panel smooth transition regression model. We assess euro area debt sustainability by analyzing the reaction of the cyclically adjusted primary balance to changes in public debt, relative to GDP. Our analysis relies on AMECO data for 18 euro area economies from 2000 to 2019. The estimation results show that there are two different regimes in the euro area: a high and a low-debt regime. These distinctions also reveal the heterogeneous behaviour of the cyclically adjusted primary balance across the distribution of the debt-to-GDP ratios. The coefficients are positive for both regimes, the coefficient in the low-debt regime, however, is not statistically significant. For a sub-sample of highly indebted countries, we find a statistically significant negative (positive) reaction coefficient for the low (high) debt regime. Thus, debt sustainability seems to be given in the high-debt regime. Several robustness tests support our findings.

The rest of this paper is organized as follows: Section 4.2 briefly discusses the theoretical background, section 4.3 introduces the PSTR methodology and section 4.4 presents the empirical outcome based on estimation using lagged debt ratio as the transition variable. In section 4.5, the concept of fiscal space is explored and subsequently used as a transition variable in the PSTR framework. Finally, section 4.6 summarises the main results.

4.2 Theoretical background

The analysis of debt sustainability by means of the fiscal reaction function studies fiscal policy decisions by estimating the response of the primary surplus to changes in the public debt relative to GDP. Indeed, the primary balance is influenced by many factors - several of them outside the control of the government/ fiscal authorities. Thus measuring (pure) discretionary fiscal policy decisions could be challenging. Therefore, in the literature, the cyclically adjusted primary balance is usually employed to proxy this effect, as also acknowledged by Golinelli and Momigliano (2006).

Based on the public budget of an economy that consists of government revenues (mainly taxes) and public spending, we could assume that a government determines the primary surplus (i.e. surplus net of interest payments) to GDP ratio, $s(t) = S(t)/Y(t)$, such that it is a positive linear function of the public debt-to-GDP ratio, $b(t) = B(t)/Y(t)$, and of a term that is independent of public debt, $\psi(t)$ (see Bohn, 1995, 1998, Greiner and Fincke, 2015, Afonso and Jalles, 2019). Thus, the primary surplus ratio can be written as

$$s(t) = \beta(\cdot)\, b(t) + \psi(t), \tag{4.1}$$

where $\beta(\cdot)$ is the reaction coefficient determining how strong the primary surplus reacts to changes in the public debt ratio and that is allowed to be variable - here being estimated as a function depending on the debt-to-GDP ratios. The parameter $\psi(t) \in \mathbb{R}$ is affected by other economic variables, like transitory government spending (such as social, climate or educational outlays etc.). As regards $\psi(t)$, we posit that it is bounded above and below by a certain finite number that is constant over time. Additionally, $\psi(t)$ cannot be completely controlled by the government: it can influence that coefficient only to a certain degree because $\psi(t)$ is also affected by the business cycle, for example, that affects the economy temporarily.

With a theoretical model, it can be shown that a strictly positive reaction coefficient on average, such that $\lim_{t\to\infty} \int_{t_0}^{t} \beta(\mu) d\mu = \infty$, implies that the debt policy of a government is sustainable, (see for instance, Greiner and Fincke (2015)). The reaction coefficient $\beta(\cdot)$ can be time-varying or a nonlinear function of public debt or even be negative for some time periods. However, on average, that coefficient must be positive. We implicitly assume that the primary surplus relative to GDP can grow without an upper bound, which could constitute a limitation. However, a positive but small reaction coefficient on average does not necessarily guarantee a bounded debt-to-GDP ratio.

Finally, we want to point out that we allowed the reaction coefficient to be varying/nonlinear. Empirical evidence from the time series context shows that the reaction coefficient could be time-varying and not always constant (see Greiner and Fincke, 2015, chap. 2.2-2.5 or Owusu et al., 2021). It should be noted, too, that a linear model with time-varying coefficients can be seen as an approximation of a nonlinear model, and the approximation is good if the parameter changes smoothly (cf. Granger, 2008). When analyzing the response of the primary surplus to variations in public debt, there is evidence that the reaction depends on the magnitude of the public debt ratio. Our analysis takes into account these considerations and detects the heterogeneity data-driven and determines the development across the distribution of the debt ratios.

4.3 Methodology

Regarding the methodology, we apply the panel smooth transition regression model according to Gonzalez et al. (2017). A two-regime PSTR is specified as

$$y_{it} = \mu_i + \beta_0 x_{it} + \beta_1 x_{it} g(q_{it}; \gamma, c) + u_{it} \tag{4.2}$$

where i is the individual within the panel whilst t is the time dimension. The variable y represents the response variable, μ represents the individual effect which is time-invariant and allowed to correlate with the regressors, x denotes the covariates which are assumed to be exogenous, and β represents the coefficients or parameters to be estimated. The function $g(q_{it}; \gamma, c)$ is the transition function which is observable, continuous and bounded. The variable q_{it} is the transition variable on which the regime-switching is conditioned, c is a vector of location parameters, whilst γ measures the slope of the transition function. The transition function is captured by a logistic regression function (see Teräsvirta (1994, 1998)), so that it is bounded between 0 and 1 as seen below

$$g(q_{it}; \gamma, c) = (1 + exp(-\gamma \prod_{j=1}^{m} (q_{it} - c_j))^{-1} \tag{4.3}$$

where $\gamma > 0$ determines the smoothness of the transition from one regime to another. The location parameter c captures the threshold between the two extreme regimes with transition functions $g(q_{it}; \gamma, c) = 0$ and $g(q_{it}; \gamma, c) = 1$. The index m determines the number of regimes and could be more than one depending on the variations in the parameter. For instance, when $m = 1$, the model is characterised by two extreme regimes associated with high and low values of the transition variable (q_{it}). In that case, the coefficients from equation 4.2 switches between β_0 and $\beta_0 + \beta_1$ and the change is centred around c_1. When $m = 2$, the transition function attains its minimum at $(c_1 + c_2)/2$ and attains its maximum at 1 for both low and high values of the transition variable. For $\gamma \to 0$, the transition function becomes a constant and the model is reduced to a linear panel fixed effects model with a homogeneous slope for any positive value of m.

Building and applying the PSTR model consists of three main procedures, namely specification, estimation and model evaluation. Model specification entails specifying a linear model with a homogeneous slope and testing the hypothesis of linearity against an alternative heterogeneous slope in the likeness of PSTR. If the PSTR model is not identified, it implies that the true data-generating process is homogeneous. The linearity test is, therefore, a significant step in the modeling process. In case homogeneity is rejected, the appropriate transition variable q_{it} and hence, the transition function $g(q_{it}; \gamma, c)$ is then determined.

Linearity/homogeneity tests can be carried out by imposing either $H_0 : \gamma = 0$ or $H_0 : \beta_1 = 0$. Due to the issue of unidentified nuisance parameters in the PSTR model, a homogeneity test requires expanding equation 4.2 by first-order Taylor expansion around $\gamma = 0$ and re-parameterising to yield

$$y_{it} = \mu_i + \beta_0^* x_{it} + \beta_1^* x_{it} q_{it} + .. + \beta_m^* x_{it} q_{it}^m + u_{it}^* \tag{4.4}$$

where $u_{it}^* = u_{it} + R_m \beta_1 x_{it}$ and r_m is the reminder of the Taylor expansion. $\beta_1^*, \ldots, \beta_m$ are the vectors of parameters. Hence, the standard linearity test implies testing for the null of $H_0^* : \beta_1^* = \ldots = \beta_m^*$ from equation 4.4 by the Lagrangian Multiplier (LM) test based on the F-distribution and Chi-square distribution. Furthermore, if the PSTR is identified, we proceed with a test to determine the optimal number of regimes also known as sequence of homogeneity test. Under this test, we refer to equation 4.4 and assume $m = 3$ so that we test for significance of a model with 1, 2 and 3 transition functions, respectively.

Specifically, we apply the LM to test the hypothesis $H_0^* : \beta_3^* = \beta_2^* = \beta_1^* = 0$, $H_{03}^* : \beta_3^* = 0$, $H_{02}^* : \beta_2^* = 0 | \beta_3^* = 0$ and finally $H_{01}^* : \beta_1^* = 0 | \beta_3^* = \beta_2^* = 0$. The null with the strongest rejection is selected as the appropriate model and hence the optimal number of regimes. See Gonzalez et al. (2017) and Teräsvirta (1994) for a detailed discussion of the test.

Regarding the second step of parameter estimation ($\phi = \beta_0, \beta_1, \gamma, c$), a combination of fixed effects procedure and Nonlinear Least Squares (NLS) is used. Firstly, the individual fixed effects in the panel are eliminated by within transformation after which the model for the transformed data is estimated by nonlinear least squares due to the nonlinearity induced by the transition function.

The selection of appropriate values of γ and c for the NLS optimisation is done by choosing starting values of the parameters and using a grid search across the parameters of the transition function such that, the parameters which yield the minimum sum of squared errors are selected. This is done via an algorithm since a manual grid search over a wide range of values could be computationally demanding. The final step of model evaluation entails a specification test also known as the test of no remaining heterogeneity to ensure the model is correctly specified. The test is a generalisation of the linearity or homogeneity test discussed above. In other words, one can test if there exists additional unmodified nonlinearity in equation (4.2). A second transition function is added to the model, and a hypothesis of its significance is tested. It entails a LM test of no extra additive linearity to ensure the final estimated model is correctly specified. In this paper, we also conduct residual diagnostics checks to ensure that our residuals are not prone to typical long panel data issues such as cross-sectional dependence.

4.4 Empirical results: debt ratio as the transition variable

Before presenting the estimation and discussing the results, we first explain the data used for our study. The data consists of 18 euro area countries with the exception of Lithuania, which was excluded because of missing data issues. Regarding the variables, we use the cyclically adjusted primary balance as the dependent variable. Our covariate of interest is the lagged debt-to-GDP ratio denoted as lagged debt ratio, i.e. b_{t-1}. Further, we control for international trade using net export as a proxy, after which we control for inflation. With motivation from Barro's (1979) tax smoothing hypothesis, we include a business cycle variable denoted as YVAR and a spending gap variable henceforth known as GVAR. YVAR is obtained as the deviation between actual GDP and its long-term trend (potential GDP). Similarly, GVAR is computed as the deviation between actual real government spending and its long-term trend. Potential GDP and long-term spending are estimated using the HP filter. As a robustness test, we estimate these variables by an alternative filter according to Hamilton (2018), and we show that the results are very similar and hence, robust.

Data for the above macroeconomic variables were obtained from the European Commission AMECO website (see AMECO (2021)). The positive effects of institutions on fiscal policy have been empirically documented (for instance, see Hallerberg and Wolff (2008), De Haan and Sturm (1994)). Hence, we control for the effect of institutions on fiscal policy using proxies, notably government effectiveness and the rule of law indicators. Government effectiveness variable captures the perception of the quality of public and civil service whilst the rule of law measures the degree to which agents conform to the rules of society, for instance, about property rights, police and the law courts. The source of the institutional variables data is the World Wide Government Indicators from the World Bank (2022). Finally, since there is also empirical evidence regarding the effect of electoral cycles on fiscal behaviour (Afonso, 2008), we therefore, construct a dummy variable to account for parliamentary elections for respective countries. We conjecture that, parliamentary elections could drive public spending. We obtained parliamentary election information from Berlin Social Science Center, PPEG (2022). Our data span a period of 20 years (from 2000 to 2019).

Table 4.1 provides insights regarding the distribution of the euro area debt ratio at the country level. Smaller eastern European countries, namely Estonia, Slovakia, Slovenia, and Finland, achieved the lowest debt ratio over the period. Conversely, high-debt exceeding 100% was recorded for some Western European countries notably Greece, Italy, Ireland, Portugal and Spain. Among them, Italy and Greece had persistently high-debt ratio with an average debt ratio of about 118% and a low interquartile range, implying that debt is consistently at a high level with

Table 4.1: Individual country summary statistics - debt ratio

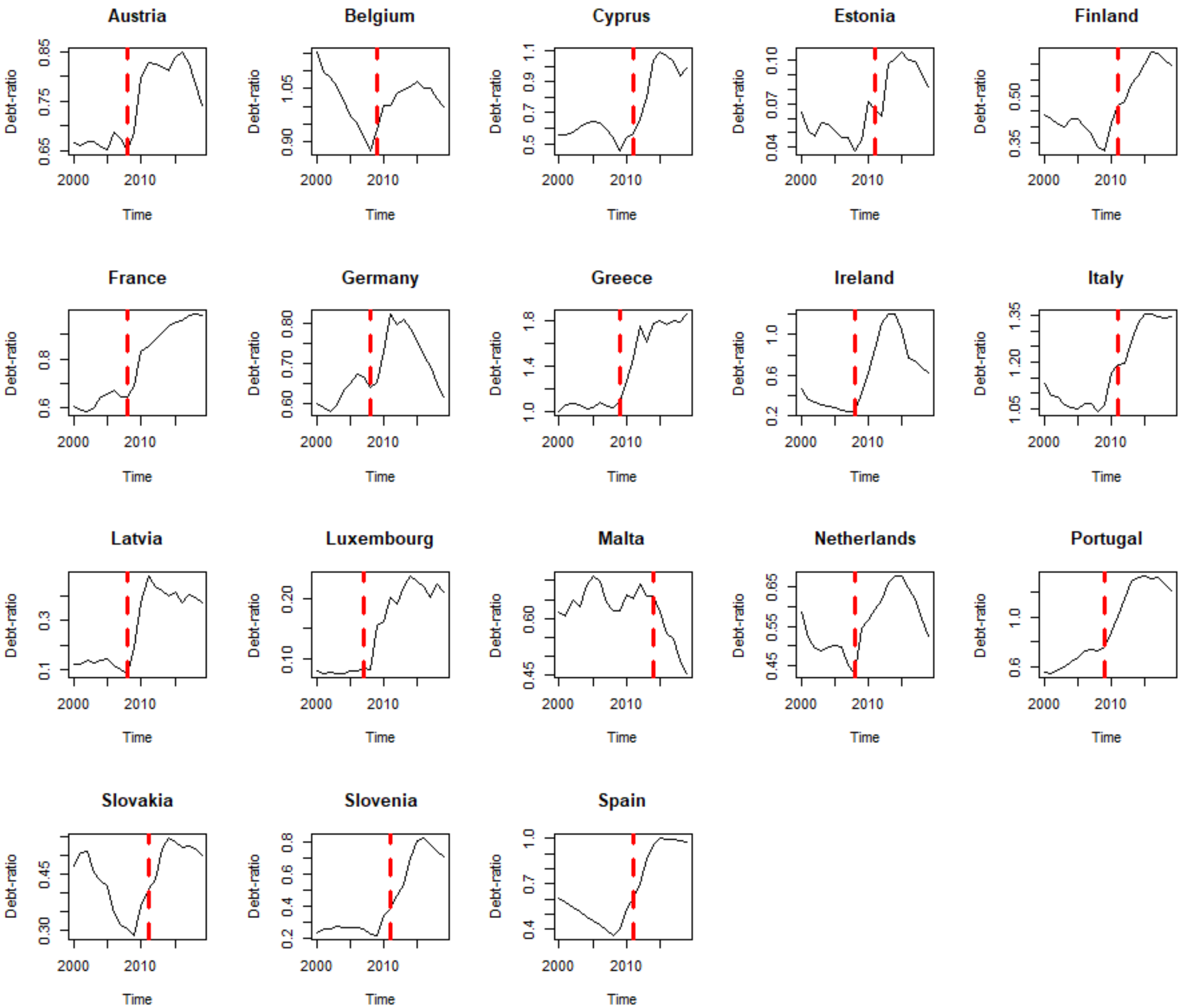
Country	Min	Pctl(25)	Median	Mean	Pctl(75)	Max	Std
Austria	0.650	0.667	0.714	0.740	0.821	0.849	0.0782
Belgium	0.873	0.991	1.03	1.02	1.05	1.15	0.0664
Cyprus	0.455	0.563	0.636	0.727	0.949	1.09	0.214
Estonia	0.0377	0.0502	0.0627	0.0689	0.0927	0.106	0.0227
Finland	0.326	0.407	0.434	0.476	0.571	0.636	0.0994
France	0.583	0.645	0.759	0.779	0.938	0.983	0.156
Germany	0.579	0.629	0.661	0.684	0.737	0.824	0.0777
Greece	0.989	1.05	1.18	1.37	1.77	1.86	0.358
Ireland	0.236	0.304	0.541	0.602	0.790	1.20	0.335
Italy	1.04	1.07	1.15	1.18	1.33	1.35	0.124
Latvia	0.0846	0.128	0.277	0.267	0.401	0.479	0.145
Luxembourg	0.0743	0.0794	0.158	0.147	0.212	0.237	0.0671
Malta	0.452	0.614	0.637	0.623	0.660	0.713	0.0677
Netherlands	0.430	0.499	0.557	0.559	0.618	0.678	0.0747
Portugal	0.542	0.663	0.817	0.929	1.27	1.33	0.308
Slovakia	0.286	0.398	0.462	0.445	0.517	0.546	0.0836
Slovenia	0.218	0.261	0.309	0.441	0.701	0.826	0.229
Spain	0.358	0.471	0.592	0.667	0.962	1.01	0.241

low variation. Therefore, it is not surprising that these countries experienced debt crisis.

Figure 4.1 provides a visual plot of the euro area debt ratio for the period. The dotted red line depicts structural breaks for each country using the test according to Zeileis et al. (2003).[1] It can be observed that debt series are characterised by two different behaviours: before the structural break and after the break. This lends some support for conditioning our regime-dependent model on the debt ratio for the euro area.

[1]The structural break test is based on F-test for potential breaks or model stability. Since the individual series are not so long, we consider 1 breakpoint. Break dates are reported in table 4.12 in the appendix.

Figure 4.1: Euro area debt-ratio with 1 structural break (break dates highlighted in dashed red).



Next, we proceed with the model specification by applying two main econometric tests discussed in the preceding section. Firstly, we conduct a homogeneity or linearity test to ascertain if the PSTR is identified. The number of regimes is subsequently determined if linearity is rejected. Our transition variable is the lagged debt-to-GDP ratio. Table 4.2 presents the results of the linearity test. The null hypothesis indicates linearity of the model. In all three models ($m = 1, m = 2$ or $m = 3$), we notice that the null is rejected at a high significance level, irrespective of whether we consider the LM test based on the chi-square distribution or F-distribution. This indicates that a nonlinear model with more than one regime is feasible.

Furthermore, we resort to the sequence of homogeneity tests to ascertain the appropriate number of regimes for the model. From table 4.3, the null H_{01}^* is the regime with the most severe rejection since its p-value is the lowest. Hence, a model with one transition (two regimes) is more suitable for the data[2].

Table 4.2: Homogeneity tests (transition variable - lagged debt ratio)

	LM_χ		LM_F	
Regimes	test	p-value	test	p-value
$m = 1$	39.37	(0.000)	7.63	(0.000)
$m = 2$	65.77	(0.000)	6.284	(0.000)
$m = 3$	75.55	(0.000)	4.743	(0.000)

Results of langrangian multiplier test of homogeneity/linearity based on chi-square (LM_χ) and F distribution (LM_F). The null hypothesis of homogeneous coefficient is tested against an alternative hypothesis of heterogeneous coefficients (PSTR).

Table 4.3: Sequence of homogeneity tests (transition variable - lagged debt ratio)

	$LM\chi$		LM_F	
m	test	p-value	test	p-value
H_{03}^*	11.97	(3.524e-02)	2.254	(4.878e-02)
H_{02}^*	29.64	(1.736e-05)	5.665	(4.909e-05)
H_{01}^*	39.37	(2.004e-07)	7.633	(8.002e-07)

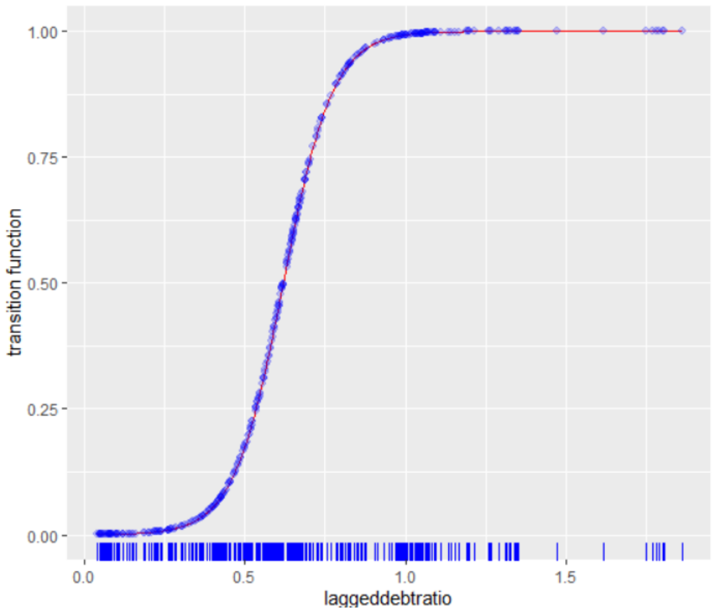
Results of LM sequence of homogeneity test based on chi-square and F distribution. The hypothesis is based on the following; $H_0^* : \beta_3^* = \beta_2^* = \beta_1^* = 0$, $H_{03}^* : \beta_3^* = 0$, $H_{02}^* : \beta_2^* = 0| = \beta_3^* = 0$ and finally $H_{01}^* : \beta_1^* = 0| = \beta_3^* = \beta_2^* = 0$

Figure 4.2 depicts a plot of the transition function based on a logistic regression function using the lagged debt-to-GDP ratio. The two extreme regions are bounded between zero and one, and the monotonic switch between the two extreme regimes where zero depicts the regime of low-debt whilst one is the extreme regime of high-debt. The transition is rather smooth between the two regimes.

In table 4.11 in the appendix, we show the distribution/characteristics of our transition function. It can be observed that there is a switch from lower to higher values between the 25th and 75th percentile, especially after 2008. Hence, the transition function depicts a time-varying behaviour. The drastic increase in the

[2]Teräsvirta (1994) presents the econometric theory behind the selection of the appropriate number of regimes.

Figure 4.2: Transition function - lagged debt ratio



average value of the transition function coincides with the global financial crisis and the European debt crisis. This is plausible since the financial crisis required some policy reaction via expansionary fiscal policies from fiscal authorities. Hence, the transition function can be characterised by two main regimes: namely, the regime before and after the crisis, which supports our earlier econometric specification test.

Next, we proceed to estimate our model characterised by two regimes. Table 4.4 presents the estimation results. It can be observed that the reaction coefficient is positive and statistically significant only in the high-debt regime. However, it is positive but not significant in the low-debt regime, indicating sustainable fiscal behaviour only for high-debt situations in the euro area[3]. This result presents evidence as we distinguish fiscal behavior in a low-debt regime and a high-debt regime in a panel data context. From a fiscal policy perspective, our finding can be interpreted in a way that in a high-debt situations the risk of running into a debt crisis is high. Hence, it is not surprising that governments react in a stabilizing manner.

Relating our result regarding the nonlinearity in the debt response directly to pre-existing literature on panel regime-switching applications could prove to be challenging, as they mainly relate to markov-switching models (stochastic approach), country studies (individual country analysis), while we apply a deterministic technique with an economic meaning of the transition variable characterized by a threshold value. Moreover, we conduct a panel data analysis with annual data and explicitly refrain from the individual country perspective allowing the economy's debt observations

[3]The implication is that, in a low-debt regime, a higher primary balance has not been generated (to offset debt) as debt-to-GDP ratio increases. Hence, sustainability of fiscal stance is not given in such a regime.

to change regimes as the debt situation changes. However, also other contributions find regime shifts with regard to fiscal policy based on markov switching approaches and often relate that to time variations or particular incidents, such as Afonso et al. (2011) for Portugal suggesting structural/ chronic fiscal problems and unsustainability. Afonso and Toffano (2013) found evidence of a sustainable situation in Germany, the UK, and Italy. Ricci-Risquete et al. (2016) found evidence of a regime shift of fiscal policy for Spain and related it to the EU accession, for instance. Aldama and Creel (2016, 2019) also identify two different regimes in fiscal policy behavior with a markov-switching method and annual data for France and the USA, each with one regime indicating sustainability and the other regime suggesting non-sustainability, pointing towards the importance of distinguishing the different situations. Accordingly, the long-run/global perspective of fiscal behavior for the US seems to be sustainable. But there are also limits regarding this comparison of the results to ours, as mentioned above.

Regarding the classical macroeconomic control variables, spending deviations $GVAR$ exert a statistically significant negative effect on the primary balance in both regimes, indicating that, expenses higher than the trend exert pressure on the budget. While inflation acts significantly negatively in high-debt situations, the statistically negative effect only shows up in the low-debt regime once we control for institutional factors (government effectiveness and rule of law). As expected, trade exerts a positive effect (except in the baseline model I in high-debt situations it is insignificant) on the primary balance. Interestingly, with regard to the fiscal policy and cyclicality the effect of the business cycle $YVAR$ varies: while in the low-debt regime, there is a positive effect, indicating counter-cyclical behaviour, for the high-debt regime, the impact on the primary balance turns out to become pro-cyclical. This (in parts) finding of pro-cyclicality is in line with several contributions in the literature on cyclicality and fiscal policy, particularly with a focus on the European context and against the background of the common regulations with the Maastricht Treaty, Stability and Growth Pact and EU membership. Some empirical results suggest pro-cyclicality prior to the regulations and a-cyclical (insignificant or neutral) afterwards, see for instance Buti et al. (1997), Gali and Perotti (2003), Wyplosz (2006). Von Hagen (2005) and Candelon et al. (2010) imply a pro-cyclicality before and after 1992. Fincke and Wolski (2016) state that the adoption of EU fiscal rules tend to change fiscal policy in a more counter-cycle direction, referring to the EU's 2004 enlargement.

Lane (2003) also finds some evidence of pro-cyclicality based on country-by-country results in OECD economies, as does Kaminsky et al. (2004) for most developing and also in middle-high income economies. However, they applied a different approach as compared to our model with a relatively large set of countries. Golinelli and Momigliano (2006) apply real-time data for euro-area economies in an OLS panel setting and find results similar to our study, e.g., a sustainable debt coefficient, counter-cyclical reaction (positive output gap coefficient) and a negative election effect. Gootjes and de Haan (2022) show with real-time data in a dynamic panel model for EU members that fiscal policy at the final/ outcome stage is pro-cyclical. Similar to our study, the debt ratio seems to behave in a sustainable manner, and political economy variables tend to have a negative effect. Larch et al. (2021) searched in a similar direction; they focused on the cyclicality of fiscal policy and find that discretionary fiscal policy tends to follow a pro-cyclical behavior. Moreover, the likelihood of that direction is enhanced strongly with high-debt ratios.

With regard to the institutional variables, our findings from table 4.4 indicate a significant positive effect of government effectiveness in both regimes, and a positive effect of rule of law in the low-debt regime and a negative effect in the high-debt regime.[4] In a high-debt situation, enforcement of the rule of law could put pressure on the budget. Furthermore, controlling for the effects of institutions led to a reduction in the magnitude of the debt coefficient (Mod II and mod IV) for both regimes. The political economic influence is proxied by the election variable, which turns out to exert a negative effect. There seems to be a loosening of the fiscal stance in the presence of elections (in line with Gootjes and de Haan (2022) and Larch et al. (2021)). The data-driven threshold value for discriminating between the high and the low-debt regime ranges between 62% and 65% depending on the model specification, which is about the reference value according to the Maastricht Treaty.

In figure 4.3, we provide a plot which explains the lagged debt contribution to the expected primary balance through the smooth transition mechanism. The graph supports a two-regime model so that in the first regime with debt less than 60%, we notice a fairly flat fit indicating a weak response of the primary balance to increment in the debt-to-GDP ratio. Conversely, a debt ratio of above 60 per cent warrants a strong response by way of increasing the primary balance which is supportive of our estimation results. The threshold value of about 62% indicates a turning point of the reaction function as can be seen in the graph.

Figure 4.7 in the appendix depicts residual diagnostics where the left panel shows a quantile-quantile plot of the residuals (model IV, table 4.4). The theoretical quantile is plotted against the sample quantile so that we are able to comment on the distribution of the residuals. We notice that most of the residual plots (data points) lie on the straight diagonal line with a few points off. One can see that the residuals are close to normal distribution. Next, the right panel shows a plot of the fitted values plotted against the residual. No particular pattern can be spotted except that the residuals are concentrated around zero and, therefore, no strong evidence of heteroskedasticity. To correct for potential serial correlation, we use feasible least

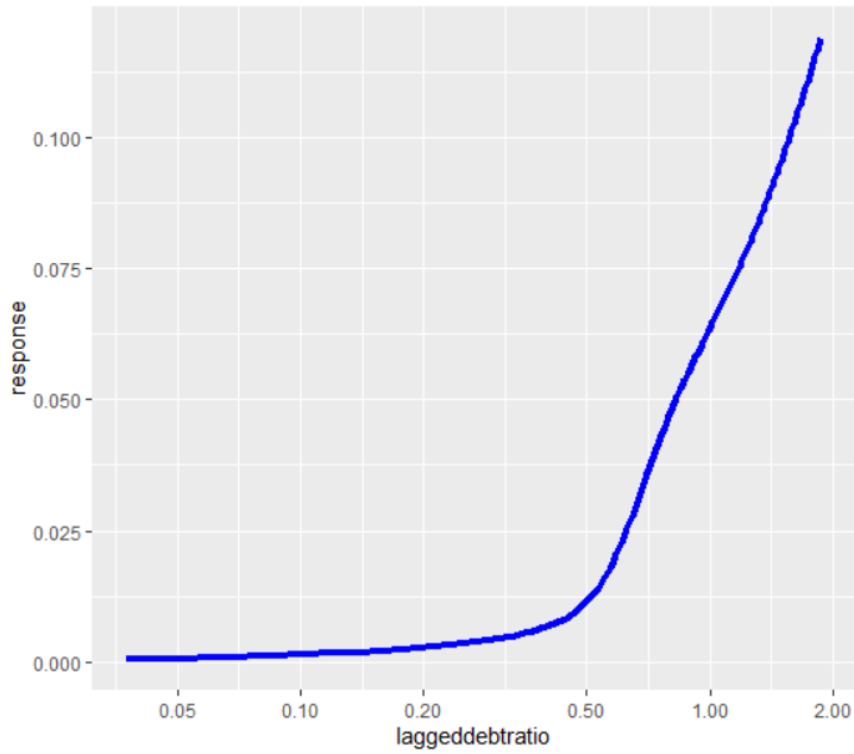
[4]Regarding the model with institutional variables (Mod II and Mod IV), the sample period is from 2002 to 2019 because of data availability. Secondly, Slovakia is excluded due to missing observations.

Table 4.4: Estimation results using PSTR - main specification

Variables	Mod I	Mod II	Mod III	Mod IV
Regime I - low-debt ratio				
lagged debt ratio	0.0144	0.0096	0.0145	0.0098
	(0.0156)	(0.0126)	(0.0139)	(0.0136)
GVAR	-0.0031***	-0.0026***	-0.0031***	-0.0026***
	(0.0002)	(0.00004)	(0.0001)	(0.000)
YVAR	0.0002***	0.0001***	0.0002***	0.0001***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Inflation	0.0029	-0.0041**	0.0035	-0.0041*
	(0.0036)	(0.0013)	(0.0034)	(0.0021)
Net export	0.0004***	0.0001***	0.0004***	0.0001***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Govt Effectiveness		0.0099*		0.0096*
		(0.0048)		(0.0051)
Rule of Law		0.0181***		0.0184***
		(0.0067)		(0.0072)
Election dummy			-0.0011***	0.0009***
			(0.0001)	(0.0000)
Regime II - high-debt ratio				
lagged debt ratio	0.0638***	0.0578***	0.0639***	0.0578***
	(0.0058)	(0.0047)	(0.0052)	(0.0044)
GVAR	-0.0007***	-0.0005***	-0.0007***	-0.0004***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
YVAR	-0.0002***	-0.0001***	-0.0002***	-0.0001***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Inflation	-0.0571***	-0.0206***	-0.0566***	-0.0195***
	(0.0022)	(0.0021)	(0.0021)	(0.0018)
Net export	0.00007	0.0001***	0.00007*	0.0001***
	(0.00004)	(0.0000)	(0.00003)	(0.0000)
Govt Effectiveness		0.0381***		0.0379***
		(0.0046)		(0.0041)
Rule of Law		-0.0410***		-0.0409***
		(0.0059)		(0.0054)
Election dummy			-0.0009***	-0.0012***
			(0.0000)	(0.0000)
Threshold value(c)	0.62	0.65	0.62	0.65
Number of observations	360	306	360	306

Estimation of $y_{it} = \mu_i + \beta_0 x_{it} + \beta_1 x_{it} g(q_{it}; \gamma, c) + u_{it}$ using PSTR. Where *, ** and *** indicate statistical significance at 1% , 5% and 10% respectively.

Figure 4.3: PSTR plot of response and lagged debt ratio



The above figure represents a plot of the response function against the lagged debt ratio. The response function is given by $[\beta_0 + \beta_1 g(q_{it}; \gamma, c)]x_{it}$.

square (procedure) based on Cochrane Orcutt approach, where an autocorrelation coefficient is estimated based on residuals and pre-multiply the resulting coefficient by the design matrices and the dependent variable leading to uncorrelated errors. Hence, we report the true standard errors (devoid of serial correlation) for all specifications.

Since the countries belong to an economic and monetary union, the likelihood of cross-sectional correlation is quite high since such countries could make similar policy decisions, for instance, responding to shocks similarly. Cross-sectional dependence affects the size properties of most econometrics test and hence, renders inferences invalid (Banerjee et al., 2004). Figure 4.8 in the appendix depicts a residual plot of the PSTR model (Mod IV in table 4.4), where we look for evidence of potential cross-sectional dependence in the country-specific residual plot. One can observe that the residuals are mostly unique, which is evidence against the strong possibility of cross-sectional dependence from our estimated PSTR model. Finally, in figure 4.9 (appendix) we compare the fitted primary balance according to the PSTR (Mod IV in table 4.4) to the actual primary balance. It can be observed that with the exception of countries like Cyprus, Latvia, Luxembourg, Slovenia and Estonia, in most cases the fitted PSTR model tracks the actual primary balance quite well.

Recall that the specification in table 4.4 depicts a panel model with only individual effect. Indeed, if time effects are significant, this could potentially lead to model misspecification. We test for the presence of individual only, time only and a combination of both individual and time effects jointly using an F-test on a simple linear specification. Results in the appendix (table 4.15) reveal that a model with only time effect is not appropriate but rather a model with individual effects or joint effect (individual and time effects) are significant. We therefore, specify a PSTR model that considers both individual and time effect shown in table 4.14 in the appendix. The result is not different from our main specification (model with only individual effect (table 4.4)) except little changes in the magnitude of some of the coefficients. Moreover, GVAR is not statistically significant in the high-debt regime. Otherwise, the story is basically the same: debt sustainability is given in the high-debt regime and a higher magnitude of reaction coefficient as compared to the lower debt regime.

4.4.1 Sensitivity analysis and robustness checks

We conduct a sensitivity analysis to ascertain how changes in the control variables (mainly the other macroeconomic variables) affect our debt sustainability parameter. From table 4.5 (model specification A), we begin by estimating a model with only lagged debt as the regressor, after which we add the other variables to the model. Results point to a lack of sustainability in the low-debt regimes, whilst fiscal sustainability is given in the high-debt regime. The same can be said about specifications B to D, which all confirm that varying the control variables does not change the main message. At low-debt levels, the reaction function is not statistically significant to infer sustainability. However, as debt crosses a particular threshold, we notice a strong response in terms of adjustment of primary balance to infer sustainability. One additional remark worth discussing is the role of inflation in the estimations. From model specifications B and C, it can be seen that controlling for inflation leads to higher debt coefficient (roughly four times the magnitude of the debt coefficient in specifications A and D). Since inflation reduces the real value of government debt, accounting for inflation leads to a higher reaction coefficient, as observed from the estimates.

We also conduct a robustness test where we consider an alternative to the HP filter used for de-trending the GDP and real government spending. We used the filter proposed by Hamilton (2018) to construct GVAR and YVAR variables and re-estimated the model. It can be observed from table 4.6 that the results are similar to our main results in table 4.4. Fiscal sustainability is not given in the first regime of low-debt, whilst the second regime of high-debt is found to be sustainable by way of a positive and statistically significant reaction coefficient of the lagged debt ratio. The control variables are similar, with only little variations. Once again, it can be observed that controlling for the effects of institutions (via government effectiveness

Table 4.5: Sensitivity Analysis

Variables	Coefficient	Std error		Coefficient	Std error
		Model A			
Regime I			**Regime II**		
lagged debt ratio	0.0112	0.0095		0.0171***	0.001
Number of observations	360			360	
Variables	Coefficient	Std error		Coefficient	Std error
		Model B			
Regime I			**Regime II**		
lagged debt ratio	0.0124	0.2391		0.06975***	0.0212
Net export	0.0003	0.0011		0.00017	0.00037
Inflation	-0.0019	0.0082		-0.0696***	0.0383
Number of observations	360			360	
Variables	Coefficient	Std error		Coefficient	Std error
		Model C			
Regime I			**Regime II**		
lagged debt ratio	0.0119	0.1461		0.0687***	0.0153
Net export	0.0003	0.0007		0.00015	0.00022
Inflation	-0.0017	0.02054		-0.0683***	0.0210
YVAR	0.00001	0.0001		-0.0001***	0.000
Number of observations	360			360	
Variables	Coefficient	Std error		Coefficient	Std error
		Model D			
Regime I			**Regime II**		
lagged debt ratio	0.0119	0.0230		0.0179***	0.0051
Net export	0.0005***	0.0003		-0.00003	0.00006
YVAR	0.0003	0.0002		-0.0002***	0.0001
GVAR	-0.0032***	0.0006		-0.0009	0.0006
Number of observations	360			360	

Estimation of $y_{it} = \mu_i + \beta_0 x_{it} + \beta_1 x_{it} g(q_{it}; \gamma, c) + u_{it}$ using PSTR. Where *, ** and *** indicate statistical significance at 1% , 5% and 10% respectively.

index and rule of law) resulted in a reduction in the magnitude of the lagged debt coefficient for both regimes. The business cycle effect (once significant) changes to mainly pro-cyclical in the low-debt regime, while it seems to be sensitive to the control variables in the high-debt situations. However, given that our main focus is on debt sustainability behaviour - the results point to the robustness of our model results to two different applied filtering methods for the cyclical variables (HP and Hamilton filter).

4.4.2 Estimation of sub-sample

In this subsection, we consider the estimation of a special group of countries in the euro area with relatively high debt-to-GDP ratios. Some of these countries run into high-debt problems, necessitating a bail-out intervention by the European Union, the European Central Bank and the IMF. They include Greece, Ireland, Italy, Portugal and, Spain, also referred to as GIIPS groups of countries. Using the same sample period, 2000-2019, we estimate a PSTR model for this group of countries.

The results from table 4.7 indicate a different pattern compared to the whole euro area: the reaction coefficient is negative and statistically significant in the first regime, indicating evidence of non-sustainable behaviour if the debt ratio is low. Thus, there is evidence that economies that found themselves in the first regime would have played a Ponzi game: Instead of raising the primary surplus in order to bear the higher debt service as public debt rises, they had raised the deficit still further. However, the reaction coefficient becomes positive and significant in high-debt situations, showing a switch towards a sustainable fiscal policy design in high-debt situations. This is also supported by the magnitude of the coefficient, compared to the whole euro area results in table 4.4 (the size of this coefficient almost doubled). The plots of the debt ratios for this sub-sample group of countries in figure 4.1 support the outcome: high and increasing debt ratios until and within the crisis and stabilisation afterwards, which could be assigned to the sustainable behaviour.

Further, transitory public spending has again a significant negative effect in both regimes, while trade has a positive effect on the primary balance. The business cycle effect appears rather small and counter-cyclical except for the specification controlling for institutional variables. The effect of inflation changes as it has a positive effect in a low-debt regime. However, it becomes strongly negative for high-debt situations for the group of GIIPS economies. These results clearly reflect the difficult situation of these economies being hit hard by the debt crisis. This also becomes visible by the threshold value, which distinguishes the low-debt regime from the high-debt regime. It has risen from 92.8% to 97.9% of GDP for this group of countries. The threshold value for Mod IV in the last column even exceeds 100%. These results align with Legrenzi and Milas (2012), also finding high-debt ratio thresholds for Italy and Greece of about 87%, but, their results refer to a historical time se-

Table 4.6: Robustness test - using alternative filter for GVAR and YVAR

Variables	Mod I	Mod II	Mod III	Mod IV
Regime I - low-debt ratio				
lagged debt ratio	0.0119	-0.0046	0.0131	-0.0044
	(0.0098)	(0.0423)	(0.0093)	(0.0094)
GVAR	0.0005***	-0.0006**	-0.0005***	-0.0005***
	(0.00004)	(0.0000)	(0.0000)	(0.0000)
YVAR	-0.0003	-0.0003***	-0.0003***	-0.0003*
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Inflation	0.0072	-0.0042	0.0073	-0.0043
	(0.0042)	(0.0093)	(0.0042)	(0.0099)
Net export	0.0002***	0.00001	0.0002***	0.00001
	(0.0000)	(0.0001)	(0.0000)	(0.0001)
Govt Effectiveness		0.0366***		0.0365***
		(0.0025)		(0.0023)
Rule of law		-0.0055		-0.0054
		(0.0128)		(0.0129)
Election dummy			-0.0006***	-0.0005
			(0.0001)	(0.0005)
Regime II - high-debt ratio				
lagged debt ratio	0.0740**	0.0641***	0.0735***	0.0637***
	(0.0126)	(0.0067)	(0.0124)	(0.0065)
GVAR	-0.0002***	-0.0002	-0.0002***	-0.0002
	(0.0000)	(0.0002)	(0.0000)	(0.0002)
YVAR	0.0002***	-0.0001	-0.0002***	-0.0001
	(0.0000)	(0.0001)	(0.0000)	(-0.0001)
Inflation	-0.0674***	-0.0300***	-0.0651***	-0.0280**
	(0.0119)	(0.0045)	(0.0115)	(0.0044)
Net export	0.0002***	0.0003***	0.0002***	0.0003***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Govt Effectiveness		0.0329***		0.0323***
		(0.0012)		(0.0014)
Rule of law		-0.0350***		-0.0347***
		(0.0012)		(0.0008)
Election dummy			-0.0053***	-0.0045***
			(0.0006)	(0.0009)
Threshold value(c)	0.651	0.655	0.646	0.655
Number of observations	306	289	306	289

Estimation of $y_{it} = \mu_i + \beta_0 x_{it} + \beta_1 x_{it} g(q_{it}; \gamma, c) + u_{it}$ using PSTR. Where *, ** and *** indicate statistical significance at 1% , 5% and 10% respectively. Potential GDP and potential spending used for computing YVAR and GVAR respectively were constructed using the filter according to Hamilton (2018).

ries and consider only individual countries. Our study explicitly focuses on panel data analysis without country-wise distinction. Government effectiveness is again significantly positive, and the rule of law negatively affects the budget. Institutional variables appear to be relevant for shaping the fiscal policy performance. Interestingly, the election dummy in the GIIPS group has a positive budget effect in low-debt situation and a negative impact in high-debt regimes. Again, under a high-debt burden, there is evidence that the budget loosens during times of elections. Certainly, Greece, one of the countries in the sample, is a special case, particularly affected by the debt crisis and experienced an extremely high debt ratio and fiscal stress as can be recognized from table 4.1 and figure 4.1. Thus, we estimated the GIIPS sub-sample again without Greece. The results can be found in table 4.13 in the appendix. That outcome supports our GIIPS finding: fiscal behaviour in low debt situations seems unsustainable while it becomes sustainable in the high debt regime. Italy, Ireland, Portugal and Spain modified their fiscal behavior towards stabilization once the debt situation exerted high pressure. The effects of the control variables change only little. This similarity of the results is not too surprising as the debt pattern seems to be quite alike, as depicted in figure 4.10.

Figure 4.4 captures the transition function for the model estimated in table 4.7 using PSTR for GIIPS countries. It can be realised that most of the distribution of the transition variable is concentrated in the high-debt regime. Additionally, we also notice that the threshold value at which the model switches from a low to a high regime is close to 100%. This is not so surprising as evidence shows these countries have run high debt-to-GDP ratios over time.

4.5 Empirical results: debt sustainability and fiscal space

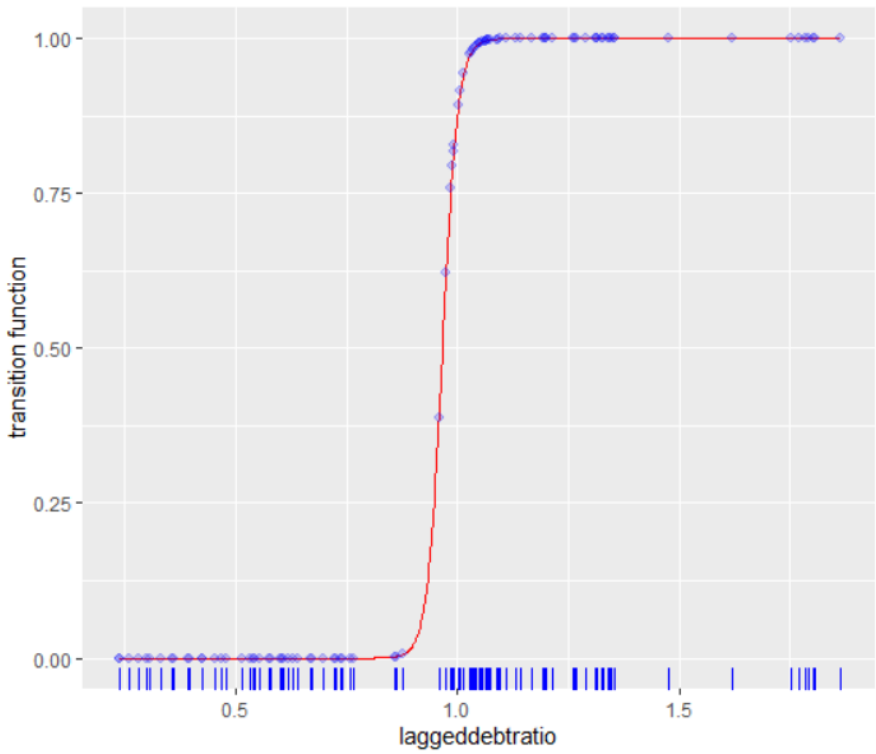
One of the issues confronting global economies has been the degree to which most countries have room for fiscal manoeuvre (Ostry et al. 2010) without necessarily putting fiscal sustainability in jeopardy. For instance, in periods of dire economic circumstance (fiscal fallout of financial crisis, pandemic or wars), the question arises whether governments have enough room to embark on expansionary fiscal policy to stimulate aggregate demand without pushing debt to unsustainable levels. Some proponents of the policy debate argues that economies with higher fiscal space should employ the available budgetary room for manoeuvre in order to stimulate economic growth, for instance (Nerlich and Reuter, 2016). This is supported by Romer and Romer (2019), who argue that countries facing crisis with adequate fiscal space are able to take much more aggressive fiscal action as compared to situations where countries have limited fiscal space. Hence, the level of a country's fiscal space is crucial to its fiscal policy actions, and more importantly sustainability of its fiscal stance.

Table 4.7: PSTR estimation results for sub-sample - GIIPS group of countries

Variables	Mod I	Mod II	Mod III	Mod IV
Regime I - low-debt ratio				
lagged debt ratio	-0.0466***	-0.1805***	-0.0426***	-0.1799***
	(0.0099)	(0.000)	(0.0008)	(0.000)
GVAR	-0.0032***	-0.0034***	-0.0032***	-0.0031***
	(0.000)	(0.000)	(0.000)	(0.000)
YVAR	0.0001**	-0.00002**	0.0001**	-0.00002**
	(0.000)	(0.000)	(0.000)	(0.000)
Inflation	0.0403***	0.2225***	0.0278***	0.1546***
	(0.0148)	(0.000)	(0.0021)	(0.000)
Net export	0.00008***	0.0007***	0.0001***	0.0008***
	(0.000)	(0.000)	(0.000)	(0.000)
Govt Effectiveness		0.0025***		0.0130***
		(0.000)		(0.000)
Rule of law		-0.1644***		-0.1623***
		(0.000)		(0.000)
Election dummy			0.0079***	0.0160***
			(0.0002)	(0.000)
Regime II - high-debt ratio				
lagged debt ratio	0.1266***	0.1224***	0.1304***	0.1171***
	(0.0039)	(0.000)	(0.0007)	(0.000)
GVAR	-0.0007***	-0.0003***	-0.0005***	-0.0003***
	(0.0002)	(0.000)	(0.000)	(0.000)
YVAR	0.0001***	0.0001***	0.0001***	0.00004***
	(0.000)	(0.000)	(0.000)	(0.000)
Inflation	-0.133***	-0.2317***	-0.1440***	-0.2652***
	(0.0134)	(0.000)	(0.0023)	(0.000)
Net export	0.0003**	0.0003***	0.0001***	0.0002***
	(0.000)	(0.000)	(0.000)	(0.000)
Govt Effectiveness		0.0367***		0.0331***
		(0.000)		(0.000)
Rule of law		-0.0425***		-0.0338***
		(0.000)		(0.000)
Election dummy			-0.0036***	-0.0067***
			(0.0002)	(0.000)
Threshold value(c)	0.965	0.928	0.979	1.004
Number of observations	100	90	100	90

Estimation of $y_{it} = \mu_i + \beta_0 x_{it} + \beta_1 x_{it} g(q_{it}; \gamma, c) + u_{it}$ using PSTR. Where *, ** and *** indicates statistical significance of 1% , 5% and 10% respectively.

Figure 4.4: Transition function for GIIPS group of countries



The aim of this subsection is to ascertain if the level of fiscal space matters for the fiscal response of the primary balance to the debt-to-GDP ratio. In other words, we want to examine the role the level of fiscal space plays in debt sustainability analysis. We, therefore estimate separate fiscal reaction functions for different regimes of fiscal space to examine fiscal policy behaviour in each regime. Notable computation of fiscal space is according to Gosh et al. (2013) and Ostry et al. (2010), who both define fiscal space as the difference between the debt limit and the current debt level. However, computing the debt limits for each country is not straightforward according to their definition as one will have to make various assumptions about the risk free interest rates, distribution of shocks to primary balance and the recovery rate in instances of default. This implies some level of complexity in the computation of fiscal space. Alternatively, we use a simple, straightforward, but intuitive approach to compute the fiscal space.

We define fiscal space as the difference between the reaction coefficient of the primary surplus to higher public debt relative to GDP and the difference between the interest rate and the GDP growth rate (interest-growth differential[5]). Hence, the fiscal space will be positive when the fiscal reaction coefficient is greater than the interest-growth differential and negative otherwise. Consider the equation below:

[5]The interest growth differential is one of the significant factors that determines the rate at which the debt-to-GDP ratio rises (Barrett, 2018)

$$\int_0^t \phi(\mu)d\mu - \int_0^t (r(\mu) - g(\mu))d\mu > 0 \tag{4.5}$$

where $r(\mu)$ denotes the real interest rate, $g(\mu)$ denote growth rate of real GDP, $\phi(\mu)$ represents the reaction coefficient of the primary balance to changes in the lagged debt ratio. From (4.5), the difference between the two expressions denotes the fiscal space. Greiner and Fincke (2015, p.9) showed that the debt-to-GDP ratio converges to a constant when the fiscal reaction coefficient exceeds the interest-growth differential as per (4.5) so that the difference between the two expressions is greater than zero. Conversely, the debt-to-GDP ratio diverges to infinity when the fiscal reaction coefficient is lower in magnitude as compared to the interest rate differential given by the equation

$$\int_0^t \phi(\mu)d\mu - \int_0^t (r(\mu) - g(\mu))d\mu < 0 \tag{4.6}$$

We call the gap (between the reaction coefficient and the interest-growth differential) the fiscal space since the bigger this magnitude is (in positive value terms), the higher the flexibility of the government in its spending options without jeopardizing sustainability of its public debt. This holds because a stronger and statistically significant fiscal reaction coefficient indicates that a stronger primary balance is generated to pay off debt as it rises. Secondly, a lower interest-growth differential implies that smaller primary surpluses are needed to stabilize the debt-to-GDP ratio. Hence, a higher fiscal reaction coefficient and a low interest-growth differential signals more room for fiscal manoeuvre since only small primary balance is needed to stabilize debt-to-GDP ratio.

4.5.1 First specification (homogeneous reaction coefficient)

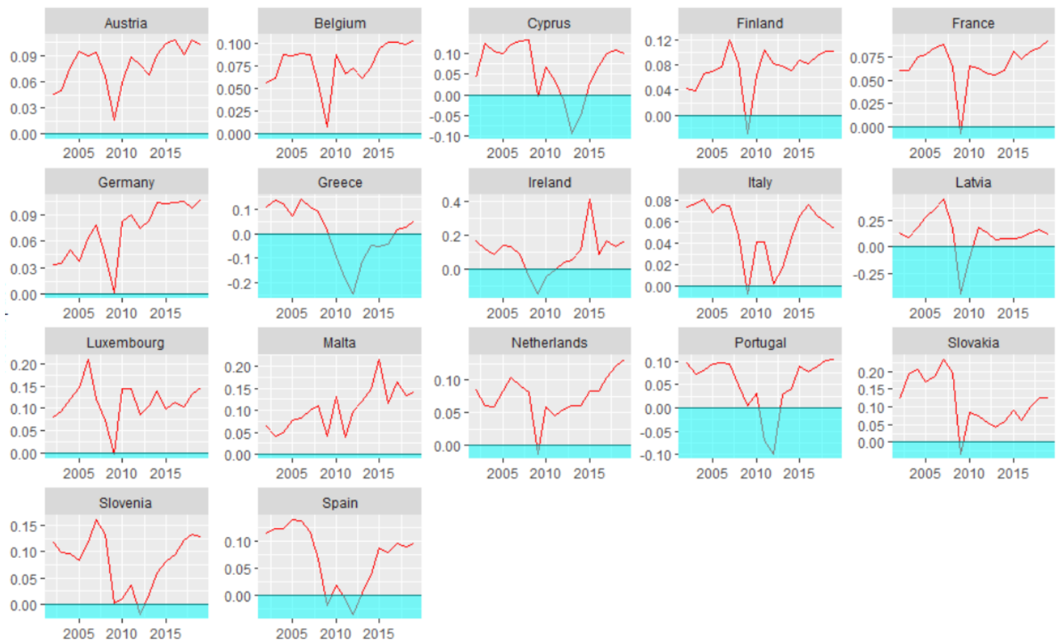
In the following, we first compute the fiscal reaction coefficient for the whole euro area based on the above definition. The assumption is that the euro area countries face one homogeneous fiscal reaction function. This can be justified by arguing that in a currency union each government commits itself to stick to some well-defined rules that limit their fiscal behavior. Hence, we combine all euro area countries in a panel and estimate a simple linear reaction function using a panel fixed effects model, without control variables and one with the deviation of government consumption (GVAR) and of output (YVAR) from their respective trends as control variables. Our main interest is to obtain a reaction coefficient of lagged debt. Using lagged debt-to-GDP ratio is justifiable because debt repayment is made in later years or periods and not necessarily in the same year. Since the data exists in annual frequency, we consider one year lag of debt-to-GDP ratio. The sample period used is from 2002 to 2019 due to data availability. Secondly, Estonia was excluded due to missing observations. Hence, we generated 306 observations consisting of 17 countries over 18 year period. The estimated coefficients are shown in (4.7) and (4.8),

$$pb_{it} = 0.056 \cdot debt_{it-1} + u_{it} \tag{4.7}$$

$$pb_{it} = 0.050 \cdot debt_{it-1} - 0.0015 \cdot GVAR_{it} + 0.0001 \cdot YVAR_{it} + u_{it} \tag{4.8}$$

Hence, we use a reaction coefficient of about 0.05 to calculate our fiscal space variable. We compute the difference between the reaction coefficient and the time-varying interest-growth differential to obtain the fiscal space for each year. The figure 4.5 depicts the fiscal space for each country, where the shaded area depicts periods of negative fiscal space. One can realize that the fiscal space is positive on average implying a bounded debt-to-GDP ratio, with the exception of Greece possibly.

Figure 4.5: Fiscal space in the euro area - first specification



In the next step, we use the fiscal space as a transition variable in a panel smooth transition regression (pstr) setting. We estimate the fiscal reaction function for the euro area for phases of low and of high fiscal space in table 4.8, where we control for other macroeconomic and institutional variables.

Table 4.8: PSTR estimation - first specification

Variables	Mod I	Mod II	Mod III	Mod IV	Mod V
Regime I - Low Fiscal Space					
lagged debt ratio	0.0197***	0.0317***	0.0168***	0.0192***	0.0164***
	(0.0004)	(0.0000)	(0.0010)	(0.0011)	(0.0005)
GVAR	-0.0002***	-0.0032***	-0.0034***	-0.0033***	-0.0032***
	(0.0001)	(0.0000)	(0.0001)	(0.002)	(0.0000)
YVAR	-0.0002***	-0.0002***	-0.0002***	-0.0001***	-0.0001
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0001)
Interest rates		-1.1250***	-2.0150***	-1.9920***	-1.743***
		(0.0036)	(0.0570)	(0.0902)	(0.0619)
Inflation			0.0065***	0.0143***	0.0170***
			(0.0051)	(0.0085)	(0.0039)
Net export				0.0124***	0.0167***
				(0.0061)	(0.0017)
Crisis Dummy					-0.0152***
					(0.0011)
Regime II - high fiscal space					
lagged debt ratio	0.0138***	0.0012***	0.0063***	0.0059***	0.0049***
	(0.0003)	(0.0000)	(0.0004)	(0.0006)	(0.0001)
GVAR	-0.0007***	-0.0006***	-0.0006***	-0.0005***	-0.0004***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
YVAR	-0.00001***	-0.00001***	0.00003***	0.00003***	0.0001***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Interest rates		-0.3803***	-0.5656***	-0.4395***	-0.4223***
		(0.0006)	(0.0117)	(0.0203)	(0.0108)
Inflation			-0.0468***	-0.0397***	-0.0378***
			(0.0019)	(0.0037)	(0.0015)
Net export				0.0098***	0.0095***
				(0.0000)	(0.0000)
Crisis Dummy					-0.0064***
					(0.0001)
Threshold value(c)	0.020	0.029	0.026	0.027	0.029
Number of observations	306	306	306	306	306

Estimation of $y_{it} = \mu_i + \beta_0 x_{it} + \beta_1 x_{it} g(q_{it}; \gamma, c) + u_{it}$ using PSTR. Where *, ** and *** indicate statistical significance at 1% , 5% and 10% respectively.

From table 4.8 we realize that public debt is sustainable in both regimes since all reaction coefficients are positive and statistically significant. However, in the low fiscal space regime the magnitude of the reaction coefficient is larger than in the high fiscal space regime. This could imply that the lack of high fiscal space induces pressure on fiscal policy to the extent that a larger primary surplus is needed to ensure a sustainable debt policy. Further, the variable GVAR has negative expected signs for both regimes implying, the transitory government spending has a negative effect on the primary balance irrespective of the regime.

As regards the output gap, one can see that in the low fiscal space regime this coefficient is negative, implying pro-cyclical behavior. However, in the high fiscal space regime, there is evidence of counter-cyclicality of fiscal policy due to a positive output gap coefficient (Mod III, Mod IV and Mod V). Moreover, the coefficient of the crisis dummy[6] in the low fiscal space regime is twice the size of the coefficient in the high fiscal space regime. The likely implication is that, in a low fiscal space regime, crisis have a stronger negative impact on fiscal policy than in a high fiscal space regime. This probably holds because countries with a higher fiscal space can react more appropriately with fiscal interventions to dampen the effects of the crisis compared to economies with a low fiscal space. Our results are in line with Jordà et al. (2016) and Romer and Romer (2019), who both found that in a crisis situation, countries with low-debt ratios do not face high economic downturns or long-lasting output losses as compared to countries with high-debt ratios because of the issue of limited fiscal space (for high-debt countries).

Trade drives fiscal surplus, as depicted by positive coefficients in both regimes. From growth regressions trade is growth-enhancing and, consequently, leads to high tax revenues and budget surpluses. Interest rates negatively affect the primary surplus in both regimes. Thus, we find empirical evidence that the disciplining effect of higher interest rate is not observable for the countries in our sample over the time span considered.

Finally, in a low fiscal space regime, inflation has a positive impact on the primary balance, whilst impacting the primary balance negatively in a high fiscal space regime. Governments with a higher fiscal space have the flexibility to increase expenditures more easily without jeopardizing sustainability and do not necessarily have to raise taxes to finance additional spending. Raising expenditures without a corresponding increase in revenues reduces the primary balance; hence, there is a negative effect of inflation on the primary balance in the high fiscal space regime.

4.5.2 Second specification (heterogeneous reaction coefficients)

In the second specification, we relax the assumption of a single fiscal reaction coefficient for the euro area. Rather, we estimate the fiscal reaction for a group of

[6]For the crisis period, we considered the period between 2008 and 2011, which coincides with the global financial crisis and the European debt crisis.

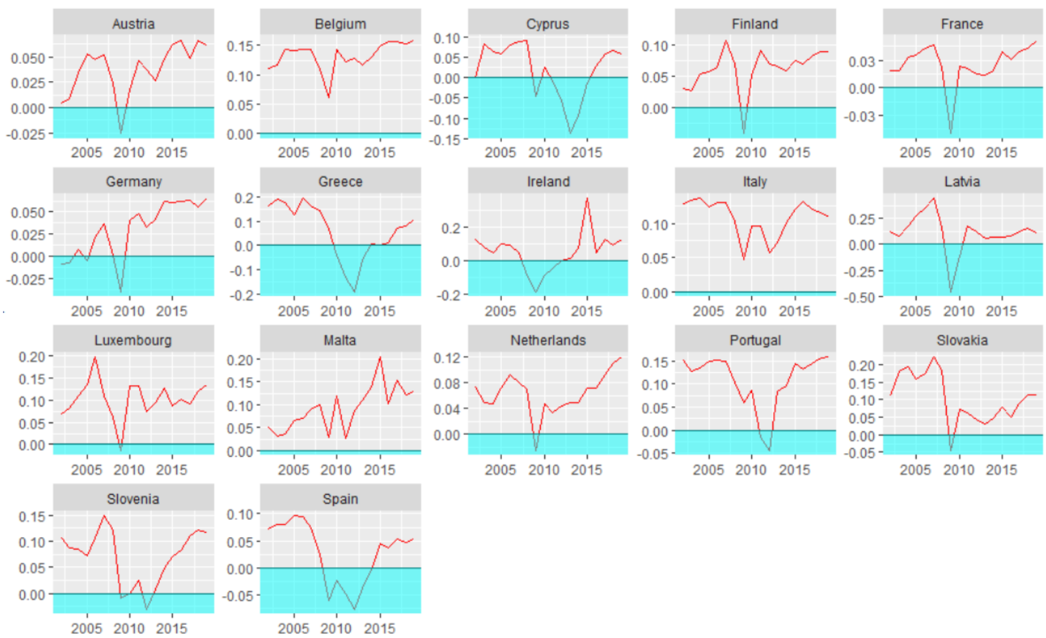
Table 4.9: Fiscal reaction coefficient based on different debt level categorization

Category	Countries	Average debt-ratio	FR coefficient
low-debt	Finland, Latvia, Luxembourg, Malta, Netherlands, Slovakia, Slovenia	0.429	0.044
Medium debt	Austria, Cyprus, France, Germany, Ireland, Spain	0.697	0.014
high-debt	Belgium, Greece, Italy, Portugal	1.143	0.111

Estimation of $pb_{it} = \beta * debt_{it} + u_{it}$ using panel fixed effects estimator to obtain the Fiscal reaction (FR) coefficient. FR coefficients represent the β for each specification namely low-debt, medium debt and high-debt.

countries based on the level of their public debt-to-GDP ratios. The assumption that all euro area countries face one reaction coefficient could be considered as too stringent given the heterogeneity within member states (for instance,heterogeneity in their debt-to-GDP ratio levels). The classification of the countries with the average debt ratio and the estimated reaction coefficients for each category, obtained by estimating a panel fixed effects model, can be found in tabel 4.9. We used the same sample size as explained in the first specification.

Figure 4.6: Plot of fiscal space for the euro area - second specification



The fiscal space for each country is determined based on the fiscal reaction function of each group as depicted in figure 4.6. Again, one can realize that the fiscal space is positive on average for the economies, possibly with the exception of Cyprus. Again, we use fiscal space as the transition variable in a PSTR framework. We estimate the fiscal reaction function for the euro area for low and high fiscal space regimes respectively (table 4.10), where we again control for other macroeconomic variables and crisis period.

Table 4.10: PSTR estimation - second specification

Variables	Mod I	Mod II	Mod III	Mod IV	Mod V
Regime I - low fiscal Space					
lagged debt ratio	0.0118***	0.0277***	-0.0055	-0.0241	0.0252***
	(0.0059)	(0.0001)	(0.0029)	(0.0205)	(0.0011)
GVAR	-0.0037	-0.0027***	-0.0038***	-0.0036***	-0.0013***
	(0.0025)	(0.0000)	(0.0006)	(0.0007)	(0.0000)
YVAR	-0.0001	-0.00003***	0.0001	-0.0008	-0.00004***
	(0.0004)	(0.0000)	(0.0001)	(0.0008)	(0.0000)
Interest rates		-0.9685***	-2.6550***	-2.6300***	-0.4425***
		(0.0021)	(0.2760)	(0.0114)	(0.0189)
Inflation			0.0358***	0.0539***	-0.0803***
			(0.0135)	(0.0137)	(0.0011)
Net export				-0.0974***	0.0013***
				(0.1373)	(0.0010)
Crisis Dummy					-0.0125***
					(0.0005)
Regime II - high fiscal space					
lagged debt ratio	0.0183***	0.0078***	0.0155***	0.0163***	0.0141***
	(0.0006)	(0.0000)	(0.0011)	(0.0006)	(0.0004)
GVAR	-0.0009	-0.0008***	-0.0008***	-0.0007***	-0.0013***
	(0.0000)	(0.0000)	(0.0003)	(0.0002)	(0.0002)
YVAR	-0.00007	-0.0001***	-0.00004	-0.00004	0.0001***
	(0.00004)	(0.0000)	(0.00003)	(0.00003)	(0.0000)
Interest rates		-0.4163***	-0.7224***	-0.6316***	0.7139***
		(0.0012)	(0.0698)	(0.0142)	(0.0142)
Inflation			-0.0610***	-0.0605***	-0.0593***
			(0.0013)	(0.0045)	(0.0011)
Net export				0.0093***	0.0230***
				(0.0027)	(0.0010)
Crisis Dummy					-0.0036***
					(0.0006)
Threshold value(c)	-0.022	0.0130	-0.029	-0.033	0.068
Number of observations	306	306	306	306	306

Estimation of $y_{it} = \mu_i + \beta_0 x_{it} + \beta_1 x_{it} g(q_{it}; \gamma, c) + u_{it}$ using PSTR. Where *, ** and *** indicate statistical significance at 1% , 5% and 10% respectively.

From table 4.10 we see that public debt is sustainable for all specifications in the high fiscal space regime. In a low fiscal space regime, the sustainability of debt is mixed. Hence, we infer that a higher fiscal space regime guarantees fiscal sustainability, whereas that does not necessarily hold in the low fiscal space regime, depending on the estimated model. However, "Mod V", which includes all the control variables (full model) implies sustainable fiscal policy. Again, the variable GVAR has a negative sign in both regimes implying that the transitory government spending has a negative effect on the primary balance. Furthermore, both the low and the high fiscal space regimes are characterized by pro-cyclical fiscal policies that can be seen by the negative coefficient of the YVAR variable.

Once again, in a low fiscal space regime, crisis have a stronger negative impact on fiscal policy than in a higher fiscal space regime. The crisis dummy coefficient in the low fiscal space regime is thrice the size of the coefficient in the high fiscal space regime. This is probably because countries with higher fiscal space can implement appropriate fiscal interventions to dampen effects of crisis more easily as compared to economies with a low fiscal space without jeopardizing sustainability. The effect of trade is the same as in the first specification analyzed in sub-section 4.5.1. Trade (net export) drives fiscal surplus as indicated by the positive coefficient in both regimes; only one specification reported a negative coefficient. Interest rates again exert a negative effect on the primary surplus in both regimes.

Finally, in a low fiscal space regime, inflation has a positive impact on the primary balance for most specifications, whilst impacting the primary balance negatively in a high fiscal space regime. Again, governments facing a higher fiscal space have the flexibility to increase expenditure without jeopardizing sustainability and do not necessarily have to raise revenues. Hence, the primary balance declines, and we obtain a negative effect of inflation on the primary balance in the high fiscal space regime. We have also seen from the above two estimations (two main specifications) that the level of fiscal space matters for the fiscal response. Particularly in a situation of low fiscal space, a higher primary balance needs to be generated as the debt ratio rises in order to guarantee fiscal sustainability. Moreover, in crisis periods, economies with higher fiscal space are able to respond to dampen the effects induced by the crisis as compared to situations of lower fiscal space.

It is important to point out that the boundedness of the public debt-to-GDP ratio is necessary to guarantee the sustainability of public debt. In particular, in a monetary union, each country should abide by the inter-temporal budget constraint since financial problems of one economy may endanger the whole union, as the Euro crisis starting in 2009 demonstrated. The primary surplus plays a vital role in assessing public debt sustainability. When the debt-GDP ratio rises, the primary surplus ratio must increase in order to ensure that the debt-to-GDP ratio is mean-reverting.

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4.6 Conclusion

This paper empirically studies debt sustainability for euro area economies based on a fiscal reaction function by applying PSTR. The PSTR method allows us to detect the existence of a threshold in the behaviour of the reaction function and refrains from the country-wise perspective. Introducing PSTR to the analysis of the fiscal response function provides an interesting perspective. It overcomes the pooling problem where one coefficient fits all and where the data is separated country-wise if one wants to estimate different regimes. PSTR allows us to distinguish debt regimes based on the data-generating process across all states of the debt ratios, independent of the country. The number of regimes is determined by the data, and the coefficients are estimated for each regime separately. In particular, they are not linked to countries but determined by the prevailing debt situations based on data points. This implies that different responses are feasible in one country according to the respective debt situations. For instance, the fiscal response turns out to be different in low-debt periods compared to periods characterised by high public debt.

The application of PSTR to euro area economies from 2000 to 2019 reveals that there are two regimes: a low-debt regime and a high-debt regime. The relationship between the primary balance ratio and public debt is not homogeneous across the distribution of the debt ratios. A positive reaction coefficient indicating sustainability has been found for the high-debt regime, while it is insignificant for the low-debt regime. The threshold value separating the low from the high-debt regime is about 60%, in accordance with the requirements of the Maastricht Treaty. Several robustness tests support these findings. Interestingly, the situation for a sub-sample of the highly indebted GIIPS economies that suffered heavily during the debt crisis is different: the low-debt regime is characterised by a statistically significant negative reaction coefficient, indicating evidence of an unsustainable debt policy, while the estimated response for the high-debt regime is statistically significant and positive suggesting sustainability of public debt.

Further, we explored the role of fiscal space in debt sustainability analysis. We first determined the fiscal space for each country in the sample, with fiscal space defined as the difference between a fiscal reaction coefficient and the interest rate - growth rate differential. The average value of the fiscal space turned out to be strictly positive, implying that the debt-to-GDP ratio remains bounded, except for Greece and Cyprus. In the next step, we estimated the reaction coefficients for different regimes of the fiscal space. When we distinguish between 3 groups of countries, high, medium, and low public debt, we obtained a mixed picture for the low fiscal space regime. However, in the regime with high fiscal space, the reaction turned out to be positive and statistically significant, independent of the estimated model. When no distinction is made, both the estimations for the low fiscal regime and those for the high fiscal regime yield a positive and statistically significant reaction of the primary surplus to higher debt, relative to GDP.

The policy implications derived from our analysis suggest that governments should carefully assess their prevailing debt situations in order to conduct fiscal policies that guarantee sustainable levels of public debt. The results also indicate that the initial debt level matters for the fiscal response and influences the relationship, this can be seen particularly in the results of the GIIPS countries. We observe a relatively high-debt ratio in the low-debt regime (with a threshold of about 100%). It implies that even larger primary balances must be generated at a very high debt level to guarantee sustainability. This is rather difficult if policy makers do not have enough fiscal space to respond to the very high-debt levels, in which case they would have to solicit for external support, for instance, from the IMF, World Bank etc., especially in situations of dire macroeconomic circumstances or external shocks.

4.7 Appendix

Table 4.11: Characteristic of the transition variable - lagged debt ratio

Year	25th Percentile	Median	75th Percentile
2000	0.0978	0.352	0.484
2001	0.0441	0.289	0.446
2002	0.0333	0.310	0.542
2003	0.0268	0.314	0.518
2004	0.0310	0.360	0.613
2005	0.0264	0.386	0.652
2006	0.0147	0.360	0.729
2007	0.0119	0.258	0.663
2008	0.00956	0.172	0.595
2009	0.0154	0.194	0.707
2010	0.0447	0.415	0.932
2011	0.129	0.531	0.956
2012	0.126	0.728	0.988
2013	0.253	0.919	0.991
2014	0.398	0.911	0.996
2015	0.446	0.930	0.997
2016	0.374	0.903	0.995
2017	0.331	0.862	0.994
2018	0.243	0.744	0.991
2019	0.186	0.640	0.992

Time varying summary statistics of the transition function $g(q_{it}; \gamma, c)$ in a PSTR. The transition variable is lagged debt-to-GDP ratio.

Figure 4.7: Residual diagnostics

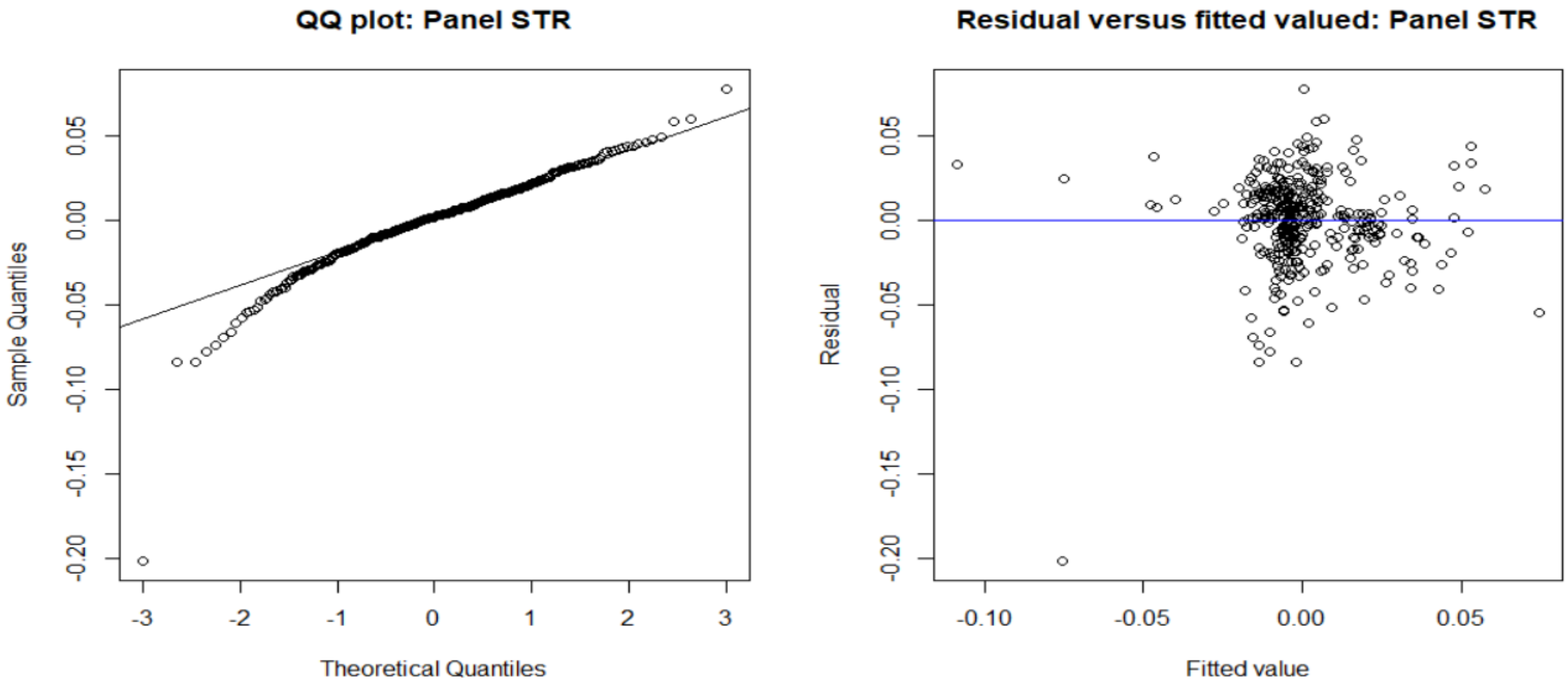

Figure 4.8: Country-specific residual plot

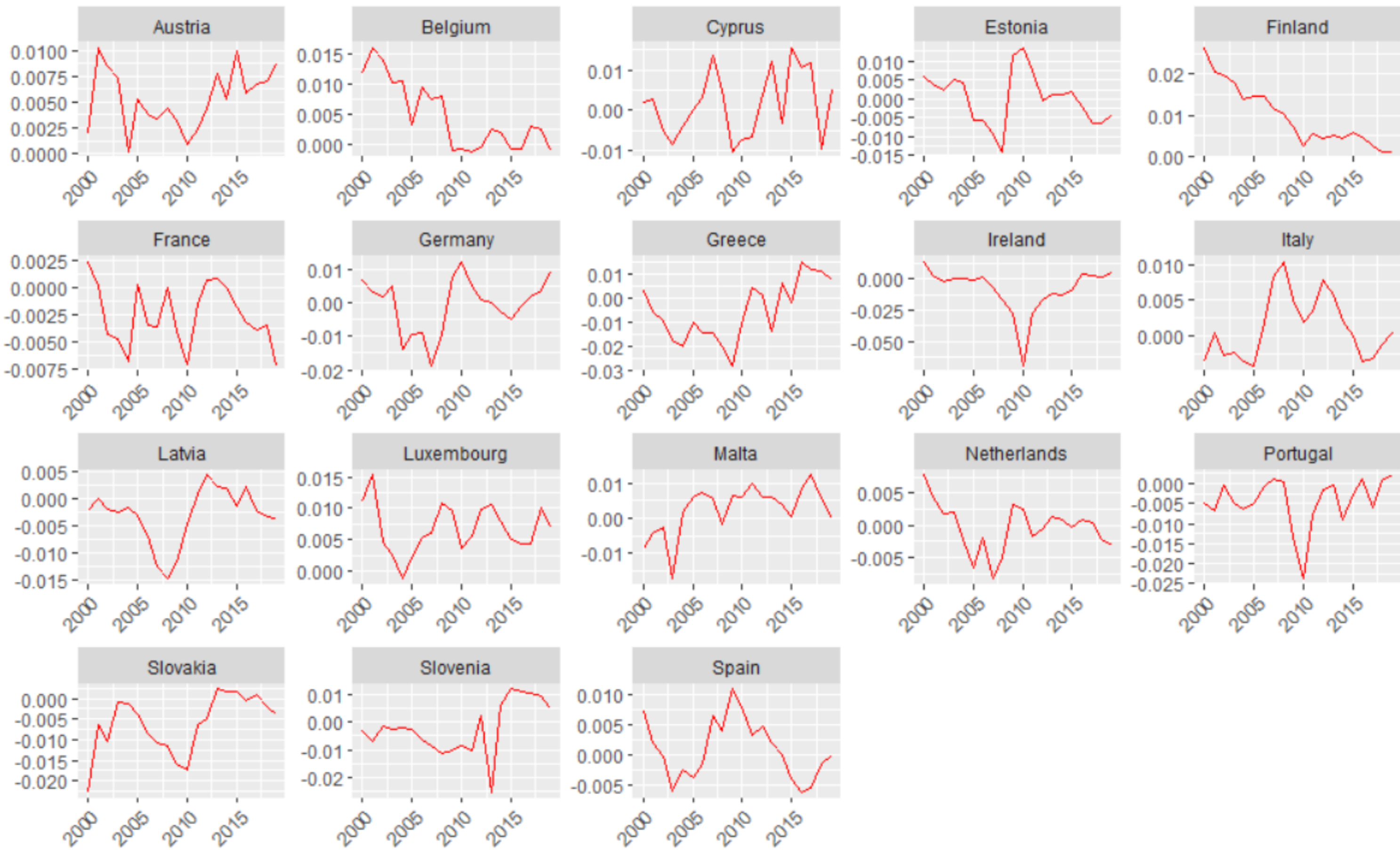

Figure 4.9: Actual and fitted primary balance

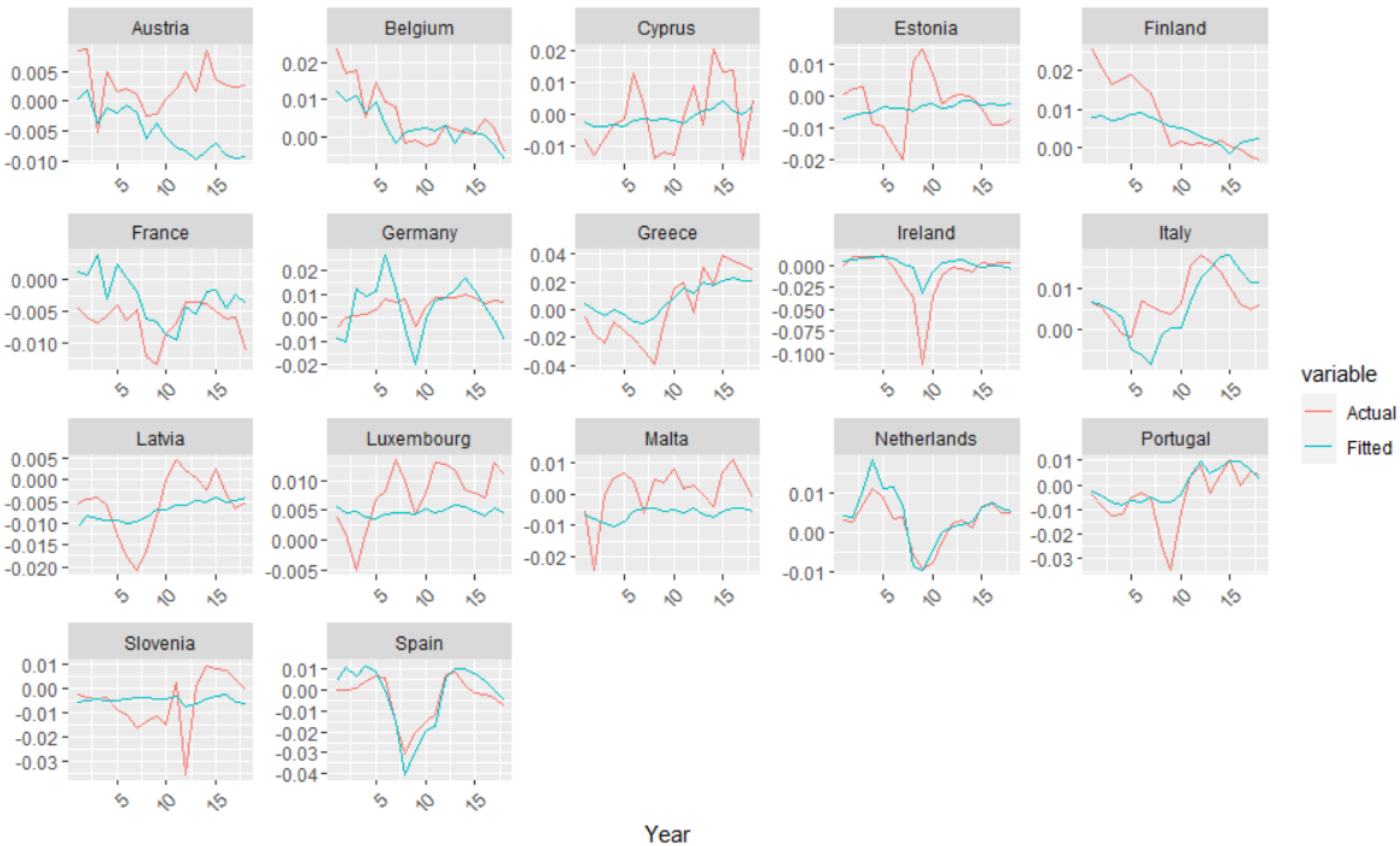

Table 4.12: Structural break dates based on test by Zeileis et al (2003) - debt-to-GDP ratio

Countries	Break date
Austria	2008
Belgium	2009
Bulgaria	2011
Cyprus	2011
Estonia	2008
Finland	2011
France	2008
Germany	2008
Greece	2009
Ireland	2008
Italy	2011
Latvia	2008
Luxembourg	2007
Malta	2014
Netherlands	2008
Portugal	2009
Slovakia	2011
Slovenia	2011
Spain	2011

Figure 4.10: Debt-to-GDP ratio plot

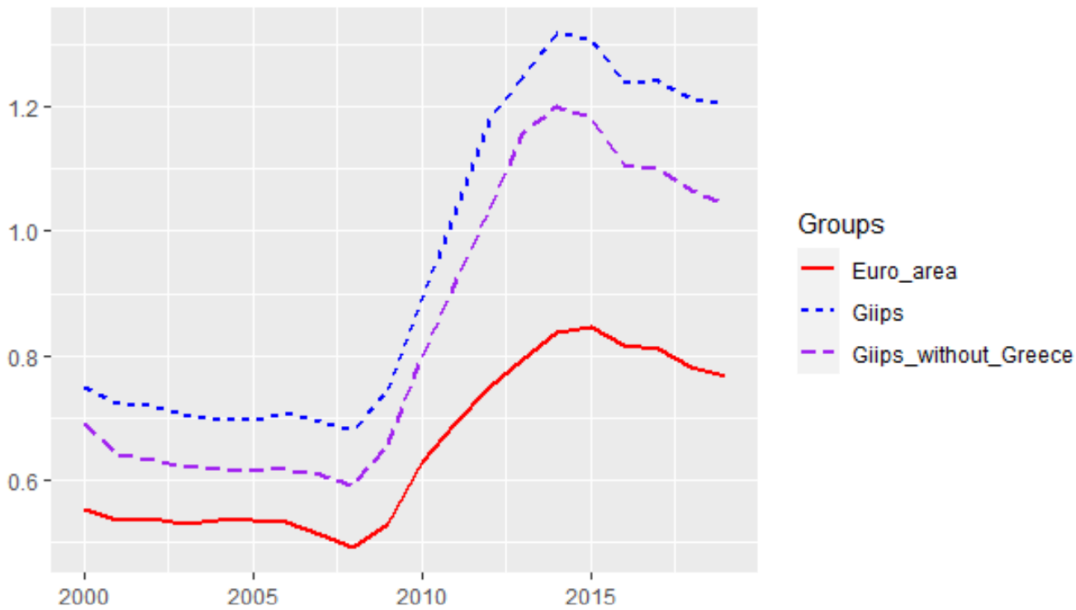

Table 4.13: PSTR estimation results for sub-sample of GIIPS (without Greece)

Variables	Mod I	Mod II	Mod III	Mod IV
Regime I - low-debt ratio				
lagged debt ratio	-0.056***	-0.1778***	-0.0567***	-0.1923***
	(0.0000)	(0.0000)	(0.000)	(0.0184)
GVAR	-0.0032***	-0.0035***	-0.0033***	-0.0034***
	(0.000)	(0.0000)	(0.0000)	(0.0000)
YVAR	0.0001	-0.0001***	0.00002***	0.00006
	(0.0001)	(0.0000)	(0.0000)	(0.00005)
Inflation	0.0661***	0.2422***	0.0614***	0.1951***
	(0.0003)	(0.0000)	(0.0001)	(0.0247)
Net export	0.00003***	0.0007***	0.0001***	0.0008***
	(0.0000)	(0.0000)	(0.0000)	(0.0001)
Govt effectiveness		0.0007***		0.0082***
		(0.0000)		(0.0017)
Rule of law		-0.1530***		-0.1645***
		(0.0000)		(0.0084)
Election dummy			0.0062***	0.0131
			(0.000)	(0.0004)
Regime II - high-debt ratio				
lagged debt ratio	0.0787***	0.0700***	0.0779***	0.0702***
	(0.0003)	(0.0000)	(0.0002)	(0.0091)
GVAR	-0.0007***	-0.0001***	-0.0001***	0.0005
	(0.0001)	(0.0000)	(0.0000)	(0.0005)
YVAR	-0.0003***	-0.0002***	-0.00002***	-0.0002***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Inflation	-0.0413***	-0.1153***	-0.0440***	-0.1689***
	(0.0002)	(0.0000)	(0.000)	(0.0291)
Net export	-0.0003***	0.0001***	-0.00003***	-0.00005***
	(0.000)	(0.0000)	(0.0000)	0.00002
Govt effectiveness		-0.0071***		-0.0175
		(0.0000)		(0.0122)
Rule of law		-0.0152***		0.0081
		(0.0000)		(0.0284)
Election dummy			0.0027	0.0032***
			(0.0001)	(0.0008)
Threshold value(c)	0.965	0.903	0.967	0.966
Number of observations	80	72	80	72

Estimation of $y_{it} = \mu_i + \lambda_t + \beta_0 x_{it} + \beta_1 x_{it} g(q_{it}; \gamma, c) + u_{it}$ using PSTR. Where *, ** and *** indicates statistical significance of 1% , 5% and 10% respectively.

Table 4.14: Estimation results PSTR - main specification (with individual and time fixed effects)

heightVariables	Mod I	Mod II	Mod III	Mod IV
Regime I - low-debt ratio				
lagged debt ratio	0.0079	0.0264	0.0080	0.0266
	(0.0047)	(0.0102)	(0.0048)	(0.0107)
GVAR	-0.0031***	-0.0028***	-0.0031***	-0.0028***
	(0.0002)	(0.0007)	(0.0002)	(0.0007)
YVAR	0.0004***	0.0004	0.0004***	0.0004
	(0.0001)	(0.0003)	(0.0001)	(0.0003)
Inflation	0.0029	-0.1202***	-0.0019	-0.1197***
	(0.0032)	(0.0361)	(0.0032)	(0.0366)
Net export	0.0005***	0.0001	0.0005***	0.0001
	(0.0001)	(0.0001)	(0.0001)	(0.0001)
Govt Effectiveness		0.0150***		0.0147***
		(0.0017)		(0.0015)
Rule of Law		0.0145***		0.0146***
		(0.0011)		(0.0013)
Election dummy			-0.0011***	0.00002
			(0.0003)	(0.0003)
Regime II - high-debt ratio				
lagged debt ratio	0.0611***	0.0573***	0.0611***	0.0574***
	(0.0035)	(0.0141)	(0.0041)	(0.00143)
GVAR	-0.0001	-0.0001	-0.0001	0.0002
	(0.0002)	(0.0008)	(0.0002)	(0.0008)
YVAR	-0.0003***	-0.0031***	-0.0003***	-0.0003***
	(0.0000)	(0.0001)	(0.0000)	(0.0001)
Inflation	-0.0761***	-0.1389***	-0.0741***	-0.1368***
	(0.0046)	(0.0095)	(0.0049)	(0.0103)
Net export	0.0000***	0.0001***	0.00005*	0.0001***
	(0.00004)	(0.0000)	(0.0000)	(0.0000)
Govt Effectiveness		0.0427***		0.0423***
		(0.0177)		(0.0177)
Rule of Law		-0.0475**		-0.0471**
		(0.0206)		(0.0205)
Election dummy			-0.0014***	-0.0012***
			(0.0002)	(0.0005)
Threshold value(c)	0.62	0.65	0.62	0.65
Number of observations	360	306	360	306

Estimation of $y_{it} = \mu_i + \lambda_t + \beta_0 x_{it} + \beta_1 x_{it} g(q_{it}; \gamma, c) + u_{it}$ using PSTR. Where *, ** and *** indicates statistical significance of 1% , 5% and 10% respectively.

Table 4.15: Test for individual and time effects - full sample

	Time effects			***Individual effects***			***Individual and time effects***		
Variables	Test-stats.	P-value	Df1 / Df2	Test-stats	P-value	Df1 / Df2	Test-stats	P-value	Df1 / Df2
Values	1.073	0.378	19 / 335	7.192	0.000	17 / 337	4.343	0.000	36 7 318
Num Obs	360			360			360		

F-test for individual and or time effects. The null hypothesis indicates the absence of significant effect (be it individual, time or both)

4.7.1 Data availability and scripts

Data used for this paper is available on my Github page (https://github.com/Benjamin-Owusu/Thesis). The data was processed and analyzed with R statistical software. Accessibility to the data and scripts will be provided by the author upon reasonable request.

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Benjamin Owusu

Lerchenstraße 65, 33607 Bielefeld, Germany
+49 (162) 476 9389
Benny.owusu@gmx.de

Education

Dec.2019 - Nov.2023 **Ph.D. Candidate in Economics** , Bielefeld University, Germany.

- **Dissertation title**: Essays on Public Debt Sustainability Analysis with Applications from Parametric and Non Parametric Panel Data Models

Oct 2017 - Oct.2019 **M.Sc. in Economic Policy**, University of Siegen, Germany.

- **Thesis title**: Estimating Monetary Policy Reaction Functions: Comparison between European Central Bank and Swedish Central Bank

Dec.2008 **Bachelor in Economics and Mathematics**, University of Ghana, Ghana.

Research Interests

- **Applied Econometrics (Panel data and time series)**
- **Applied Macroeconomics**
- **Development economics**

Academic publications

- Owusu, B.K., Boekemeier, B. and Greiner, A.(2021), "Regime-based debt sustainability Analysis: Evidence from Euro area economies", *European Journal of Political Economy, https://doi.org/10.1016/j.ejpoleco.2023.102458*
- Owusu, B.K., Bökemeier, B. and Greiner, A. (2022) "Assessing heterogeneity and non-linearity in debt sustainabilty analysis: A panel spline approach", *Empirical Economics*, vol. 64, pp. 1315–134
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- Greiner, A. and Owusu, B.K.(2020), "How to spend 750 billion euro? Applying sacrifice theory to determine Covid- 19 compensations in the EU", *Economics Bulletin*, vol.40(3), pp. 2457-2470.

Conference Presentations / Talks

- 5th INFER / ERMEES Macroeconomics Workshop 2023, University of Strasbourg, France, 1 December 2023
- 18th Warsaw International Economic Meeting, University of Warsaw, Warsaw, Poland, 4-5 July 2023
- 10th UECE Conference on Economic and Financial Adjustments. ISEG/UL – Lisbon School of Economics and Management, Lisbon, Portugal, July 2022
- 9th UECE Conference on Economic and Financial Adjustments. ISEG/UL – Lisbon School of Economics and Management, Lisbon, Portugal, July 2021

Work Experiences

Mar.2020–Current **Research Assistant** , Bielefeld University, Chair of Economic Policy, Germany.

Jun.2021–Sept. 2021 **Data Science Intern**, Matchmetrics GmbH, Bielefeld, Germany.

May.2014–Aug.2017 **Supply Chain Analyst**, TechnipFMC Energy services, Ghana.

Sep.2009–April.2014 **Supply Chain Analyst**, Halliburton Oilfield Services, Ghana.

Teaching

- MA Economic Policy Practical Tutorial (Public debt Sustainability) Winter Semester 2023/2024, Bielefeld University
- MA Economic Policy Practical Tutorial (Fiscal -Monetary Interaction) Summer Semester 2023, Bielefeld University
- MA Economic Policy Practical Tutorial (Inflation and Economic Growth)Winter Semester 2022/2023, Bielefeld University

Research Grants

- European Commission (DG ECFIN) €5000, Project: "European Governments' Fiscal Behaviour and Public Debt Holders: What Is the Financial Connection", (With Peter Claeys, Bettina Bökemeier, Andreea Stoain)

Skills

- **Computer /Software Skills:**
  R, Python, SQL, Matlab, STATA.

.

Other activities

- **Reviewer/Referee**:
  International Economics (Elsevier), Journal of Economic Integration, Economic Bulletin (ISSN 1545-2921)

Websites, Links,other

- **Google Scholar** (Link)
- **Github (https://github.com/Benjamin-Owusu)** Link
- **ORCID ID : 000-0001-7796-6514** Link

References

- **Prof. Dr. Alfred Greiner**
  Chair of Economic Policy
  Faculty of Economics and Business Admin
  Bielefeld University, Germany
  agreiner@uni-bielefeld.de

- **Prof. Dr. Dietmar Bauer**
  Chair of Econometrics
  Faculty of Economics and Business Admin
  Bielefeld University, Germany
  dietmar.bauer@uni-bielefeld.de

Bielefeld, 20th February, 2024.