Revisiting Fiscal Sustainability Analysis in the Presence of Cross-Sectional Dependence and Heterogeneity

By

Abdulkadir Bulut

4242763

Department of Economics at the University of Nottingham

This dissertation is presented in part fulfilment of the requirement for the completion of a MSc in Economic Development and Policy Analysis in the School of Economics, University of Nottingham. The work is the sole responsibility of the candidate.

Supervisor: Asst. Prof. Atsuyoshi Morozumi

December 2016

Word count: 14,560

ABSTRACT

This paper assesses public finance sustainability based on an unbalanced panel of 55 countries consisting of 31 non-advanced and 24 advanced economies over 1950-2011. We perform more recent and flexible empirical techniques following Pesaran's (2006) Common Correlated Effects Mean Group Estimator (CCEMG). The purpose of utilising this model setup that diverges from the standard panel estimators is two-fold. First, this setup addresses the possible cross-section components of the residuals, which ultimately causes endogeneity problem. Second, it accounts for the potential cross-country parameter heterogeneity in the impact of observables and unobservables. In the second part of the analysis, we examine the time evolution of the "global" fiscal reaction function. We find that while slope homogeneity assumption causes underestimation of the fiscal response function, cross-sectional correlation results in overestimation in the FE estimator.

ACKNOWLEDMENTS

Foremost, I would like to sincerely thank my supervisor Atsuyoshi Morozumi for the continuous support and guidance throughout my master thesis, for his motivation and patience.

Besides my supervisor, I would like thank to the Turkish Ministry of Education for providing financial help.

The last but not the least, I would like to dedicate this work to my cousin, my best friend, my brother Firat Bulut, who lost his life heroically in a duty, protecting civilians from terrorists in a recent coup attempt in Turkey. He has been my biggest role model since my childhood, and shaped my outlook of the world. His memory will always live with me.

Contents

LIST OF TABLES

LIST OF FIGURES

1. INTRODUCTION

Public debt sustainability has always been very significant for the current policy debate. Therefore, the risks linked with public debt levels have always been an interest for the policymakers as well as economists. In the 1970s and 1980s the debt level began to rise fast which pushed some countries to debt crises. Approximately ten years later, the history repeated itself with the uplift of public debt, which then affected the economic performance of some countries. This could simply lead to future fiscal crises. There are two ways how a high public level could affect economic activity in a negative way: Firstly, high amounts of taxes are claimed to pay back debts. This puts suppression on as well as private and government investments. Secondly, the government has to stop its spending and/or raise its taxes in order to compensate for the deficit. This could lead to a debt crisis where the government has to inflate the debt away. Both of these scenarios include economic and welfare expenses. Fiscal sustainability usually refers to a government's fiscal policies and it should be adaptable to future scenarios without any amendments. With a sustainable fiscal policy, a government should be capable of to deal with its debts in the current day as well as in the future.

There are many countries that pop into mind when the topic is about problems caused by fiscal policies. Greece leads these example countries, following by Italy, Belgium, Spain and Japan. Many solutions have already been found to deal with excessive debt accumulation and the Maastricht Treaty from 1992 is a supreme example. This treaty set out two requirements in order to partake in the Eurozone, which are the annual general government deficit at market prices must not exceed 3% at the end of the previous fiscal year, and the ratio of gross government debt-to-GDP at market prices must not exceed 60%.
Although there is no unified practical way to measure fiscal sustainability, fiscal reaction function approach has been gaining broader acceptance after the seminal work of Bohn (1998). It measures to which extent policy makers respond to rising debt level that improves primary balance. An improvement in primary balance implies that governments take measures by either reducing expenditures or increasing revenues to offset the changes in public debt.

Numerous studies have applied Bohn's fiscal reaction function (FRF) to cross country analysis. As it concerns econometric methodology, potentially important objections, largely ignored in the previous studies in FRF literature, are the implicit assumptions of cross-sectional independence (although they attempt to account for it, they do not formally test) and parameter homogeneity across countries. However, there is a growing consensus that panel data models often inherit cross-sectional dependence and parameter heterogeneity.

Cross-sectional dependence may emerge for a number of reasons such as spill over or spatial effects; or due to unobserved common components that influence all countries but most likely to a different extent, and ultimately enter into the error term (Coakley et al., 2006; Hoyos and Sarafidis, 2006 Pesaran and Tosetti, 2011). In the presence of various global shocks, the most recent financial crisis originated in the USA, addressing dependence is important, as these shocks often lead to synchronized fiscal responses across countries in traditional panel econometric methods.

Empirical exercises in FRF literature also heavily rely on estimation methods (i.e., fixed effect models) which usually allow countries to have only different country-specific fixed effects (varying intercepts), but impose an implicit assumption of slope homogeneity across units. This is the simplest but very restrictive attempt to account for heterogeneity in a panel. However, when one studies complex economic behaviours such as countries, assuming slope parameter homogeneity across sections appears to be inappropriate (Durlauf et al., 2001, Brock and Durlauf, 2001 and Eberhardt and Teal, 2011).

In this study, we first aim to tackle these two important factors which are largely ignored in the earlier literature. Our theoretical model specifications closely follow the recent work of Mauro et al. (2015) which is motivated by Bohn's (1998) argument. Differently from Mauro et al., what we add to their study is the use of more recent and flexible empirical techniques following Pesaran's (2006) Common Correlated Effects Mean Group Estimator (CCEMG) method. The purpose of utilising this model setup that diverges from the standard panel estimators is two-fold. First, it accounts for the potential cross-country parameter heterogeneity in the impact of observables and un-observables. Second, this setup addresses the possible cross-section components of the residuals, which ultimately causes endogeneity problem. As a second contribution, we take a closer look at the time evolution of the "global" fiscal reaction function. The time variation in fiscal response function has also attracted a great deal of attention in academic literature. Yet, the interest was primarily limited to single country analyses. This, in part, may be owing to the obvious concerns about panel estimators. Using panels; however, can give more valuable insights as global changes in fiscal solvency can potentially guide policy makers in individual countries to take adequate preventive measures. This is particularly relevant in light of ever-increasing globalization and complex interlinkages among world economies. We also elaborate on potential "biases" of using the FE. Results of our simulations suggest that slope homogeneity assumption causes overestimation of the FRF with the FE, while cross-sectional dependence leads to underestimation relative to the CCEMG, Thus, the time evolution of the FRF is reverse "U" shaped with the FE.

The remainder of the paper is organized as follows: The next chapter provides a summary of the relevant (but not extensive) literature. Section III discusses econometric issues encountered in the earlier literature related to cross-sectional dependence and parameter heterogeneity. Section IV summarizes

the data, model specification and econometric strategy applied. The estimation results and time variation in fiscal response functions are presented in section V. Finally, section VI concludes with a brief summary.

2. LITERATURE REVIEW

Despite the few controversies[1], there is a general agreement on the definition of fiscal sustainability in the literature. Fiscal sustainability generally refers to the capability of a public entity to maintain its current fiscal policies (i.e. expenditures, revenues and other policies) in the long period without risking its solvency. In other words, the expected present discounted value of all future primary surpluses meets the current public debt stock; or the so-called intertemporal budget constraint of a government is met. If the condition is satisfied, the public entity firstly avoids over-accumulation of the debt, secondly it can pay its debts without necessarily issuing new debt on a routine basis (No-Ponzi condition), and thirdly it bears no risk of insolvency (Aleš Krejdl, 2006).

However, there is no consensus on how to measure the fiscal sustainability empirically. The present literature has mainly followed two different methods. While the first approach is related to seeking the appropriate sustainability indicators (see, for example, Buiter, 1985; Blanchard, 1990; Uctum and Wickens, 2000; Croce and Juan-Ramón, 2003; Polito and Wickens, 2005), the second approach undertakes sustainability tests that rest on the intertemporal government budget constraints. The empirical assessments following the second method can be categorised into three approaches. The first approach tests for the presence of unit roots in public debt and budget balance series (Hamilton and Flavin, 1986; Kremers, 1988; Wilcox, 1989). The second examines the co-integration between tax revenues and government spending (Trehan and Walsh, 1988; Hakkio and Rush, 1991; Ahmed and Rogers, 1995; Haug, 1995; Quintos, 1995; Bravo and Silvestre, 2002). The third investigates whether fiscal policy reacts to increases in debt-to-GDP by increasing budget surplus (Bohn, 1998; Bohn 2005; Bohn, 2007). For this exercise, we closely follow the third approach relying on Bohn's arguments, as it will be detailed and motivated in the later section.

Since the large part of the literature concerns the definition of the public finance sustainability based on IBC, as a starting point we will now set up the derivations of the IBC, to introduce theoretical aspects of fiscal sustainability before introducing empirical specifications. This part also serves to motivate model based sustainability analysis we use; as Bohn's concerns on inapplicability of standard time series approaches (unit root and cointegration tests) employed by earlier literature (Hamilton and Flavin

[1] Pasinetti (2000) argues if the final debt ratio should be used as a result of the concept.

(1986, 1991) largely based on intertemporal budget constraint. Firstly, the focus will be on the one single period government budget constraint:

2.1. Single period budget constraint

The single period budget constraint can be expressed as follows:

$$G_t + (1 + i_t)B_{t-1} = B_t + T_t \quad (1)$$

Where G_t is the nominal government spending (exclusive of interest) in period *t*; B_t is the stock of public debt at the end of period *t*; i_t is the one period nominal interest rate on the debt on period *t*-1; T_t is the nominal tax yield. The left hand side in the equation expresses the total expenditure of the government and the right hand side indicates the sources of the government income. Equation (1) can be rewritten as follows:

$$B_t = (1 + i_t)B_{t-1} + FD_t \quad (2)$$

Where FD_t is primary (non-interest) fiscal deficit ($FD_t = G_t - T_t = -FS_t$).

2.2. Inter-temporal budget constraint (IBC) in real terms

We assume that the real interest rate is time invariant and positive: $r_{t+i} = r$ where *i*=1, 2,...,*n*. To obtain the inter-temporal budget constraint (IBC), we rewrite the one period budget constraint (1) in real terms and in conditional expectation. Based on this, solving the budget constraint recursively forward for *t+1*, *t+2*..,,∞ gives:

$$B_t = (1 + \mathrm{r})B_{t-1} - FS_t$$
$$B_{t+1} = (1 + \mathrm{r})B_t - FS_{t+1}$$
$$B_t = (1 + r)^{-1}B_{t+1} + (1 + r)^{-1}FS_{t+1}$$
$$B_t = (1 + r)^{-2}B_{t+2} + (1 + r)^{-2}FS_{t+2} + (1 + r)^{-1}FS_{t+1}$$
$$\vdots$$

$$B_t = \sum_{n=1}^{\infty} \beta^n E_t FS_{t+n} + \beta^n E_t B_{t+n} \quad (3)$$

Where β=1/(1 + r) is a discount factor, E_t is an expectation operator at time t and $FS_t = T_t - G_t$ is primary fiscal surplus. Equation (3) refers to that the initial stock of government debt B_t must offset the sum of the present values of the expected future primary balances and the future government debts from time *t* to *t+i*. However, this works if and only if the present values of the expected future government debts converge to zero when *n* goes to infinity, namely, no-Ponzi condition or equivalently transversality condition should be satisfied. To illustrate, we take the limits as *n* goes to infinity:

$$B_t = \sum_{n=1}^{\infty} \beta^n E_t FS_{t+n} + \lim_{n\to\infty} \beta^n E_t B_{t+n} \quad (4)$$

$$\lim_{n\to\infty} \beta^n E_t B_{t+n} = 0 \ \ (NPC) \quad (5)$$

With the Ponzi condition i.e. $\lim_{n\to\infty} \beta^n E_t B_{t+n} > 0$, the government funds its debts through issuing new debts on a routine basis, which implies that there will be some amount of debts that are not paid by the public entity to the debt lenders. This causes lenders to be skeptical about the government's ability to repay its debt and the credibility of the government to diminish for future debts. In this case, investors do not prefer buying government bonds and the debt will ultimately explode and cannot be maintained perpetually. However, no-Ponzi condition or transversality condition (5) does not allow the action of the government's funding through rolling its debt on a regular basis. When no-Ponzi condition is satisfied i.e $\lim_{n\to\infty} \beta^n E_t B_{t+n} = 0 \ \ (NPC)$, equation (4) becomes:

$$B_t = \sum_{n=1}^{\infty} \beta^n E_t FS_{t+n} \quad (6)$$

Equation (6) demonstrates the IBC of a government implying that the present debt should correspond to the present values of the expected primary fiscal surpluses.

2.3. Inter-temporal budget constraint with growth

It is more useful to scale the relevant variables in IBC by GDP ratios since it gives more information about a growing economy. Assuming a positive time-invariant interest and growth rate ($r_{t+i} = r$ and $\gamma_{t+i} = \gamma$, respectively where *i*=1, 2,...,*n*), we can rewrite the inter-temporal budget constraint in GDP measures:

$$b_t = \sum_{n=1}^{\infty} \theta^n E_t fs_{t+n} + \lim_{n\to\infty} \theta^n E_t b_{t+n} \quad (7)$$

Where $fs_t = (T_t - G_t)/GDP_t$ is the primary surplus-to-GDP ratio, $b_t = B_t/GDP_t$ current debt-to-GDP ratio and θ=$(1+\gamma)/(1+r)$ is discount factor. Equation (7) turns into equation (9) which is the IBC only when transversality condition or no-Ponzi condition (eq. 8) is satisfied:

$$\lim_{n\to\infty} \theta^n E_t b_{t+n} = 0 \text{ (NPC)} \quad (8)$$

$$b_t = \sum_{n=1}^{\infty} \theta^n E_t fs_{t+n} \quad (9)$$

The NPC (8) implies that the present value of the future public debt in GDP percentage should be zero while n goes to infinity in limit. Equation (9) suggests for the IBC to hold, the discounted present value of the future primary balances-to-GDP ratio must be equivalent to the initial stock of government debt-to-GDP for a sustainable fiscal policy. As can be seen at the above equations (θ=$(1+\gamma)/(1+r)$), debt ratio also depends on the divergence between the real interest rate and the average growth rate. When the debt-to-GDP ratio is positive ($b_t > 0$), the real interest rate should be higher than the growth rate ($r > \gamma$) for the sustainability condition to be granted because of the following two reasons, namely, the stabilization of the future debt and convergence to zero in limit. Hereby, positive future primary surpluses ($fs_{t+n} > 0$) on average are sufficient to ensure fiscal sustainability.

2.4. Empirical tests on fiscal sustainability

A widely used practical way to assess fiscal policy sustainability in the empirical literature is to analyse the time series properties of past fiscal data. The focus of some earlier research has been on checking the univariate properties of the stock of debt (B_t) and deficit series (FD_t) (Hamilton and Flavin[2], 1986; Kremers, 1988; Wilcox, 1989; Baglioni and Cherubini, 1993; Caporale, 1995; Uctum and Wickens, 2000). To do this, standard stationary tests such as ADF and PP have been employed. In this regard, a difference-stationary debt (i.e. B_t ~ I (0)) suggests that the relevant fiscal series is zero mean reverting

[2] The first attempt by econometric means to test sustainability is Hamilton and Flavin (1986). Their findings suggest that the hypothesis of a unit root is rejected for both real deficit and debt stock in US data. Following, other attempts present a more mixed puzzle depending on the assumptions made. For example, employing the same dataset, Wilcox (1989) suggests that when relaxing the assumption of constant interest rate and considering stochastic violations to the IBC, the IBC condition does not hold in the data.

over time, thus concluding that a government's debt policy is sustainable. In contrast, in the presence of unit root in the debt series, the no-Ponzi scheme is violated ($\lim_{n\to\infty} \beta^n E_t B_{t+n} > 0$); thus, it has been concluded that IBC does not hold, which implies that the debt grows without bound leading fiscal unsustainability.

Other empirical literature has made use of co-integration techniques to assess fiscal sustainability (Trehan and Walsh, 1991; Hakkio and Rush, 1991; Smith and Zin, 1991; Ahmed and Rogers, 1995; Quintos, 1995; Haug, 1995; Payne, 1997; Martin, 2000; Bravo and Silvestre, 2002). In this technique, the implicit hypothesis is also stationarity regarding the real interest rate[3] (Afonso, 2005). If total government spending inclusive of interest payments (G_t^o) and total revenue (T_t) are both I (1) process, but the linear combination of the series is I (0) with a co-integration vector of (1,-1), the government satisfies the IBC. This also implies that even if the public debt consistently grows, policy can be sustainable through the government's reactions to increases in debt by adjusting its spending and revenues. In other words, total spending and revenues do not diverge in the equilibrium.

However, in an influential set of studies, Bohn (1995, 1998 and 2007) questions the validity of the theoretical and the empirical framework for fiscal sustainability analysis. According to Bohn's theoretical discussion, the NPC or TC (equation) is not a sufficient condition for sustainability analysis in a stochastic environment. He argues that the future debt, revenues and expenditures in IBC cannot be formulated in terms of expected fiscal variables discounted at a safe interest rate. By making such strong assumptions, one disregards the probability distribution of the variables across states of nature. When the discounting factor is incorrectly expressed (in terms of safe interest rate), although the government policies followed a sustainable debt-to-GDP ratio, it would be labelled as unsustainable or vice versa. For these reasons, we do not follow no-Ponzi Game or equivalently expressed transversality condition (8) as a theoretical framework for empirical assessment.

Empirically, Bohn (2007) argues that the standard stationarity and co-integration tests do not work efficiently for rejecting the consistency of the past fiscal data with IBC. He proves[4] that when the debt variable (B_t) becomes stationary for any finite number differencing operations, then no-Ponzi game (equation) and B_t, T_t and G_t^o meet the IBC. The IBC is also proven to be satisfied in the case that revenues and total expenditures (inclusive of interest payment) are difference-stationary with an

[3] Assuming that the real interest rate (r_t) is stationary with mean *r*, the IBC can also be expressed as :
$\Delta B_t = G_t + (r)B_{t-1} - T_t = G_t^o - T_t$ where G_t^o is total expenditure inclusive of interest payment

$$\Delta B_t = G_t^o - T_t = \sum_{n=1}^{\infty} \beta^n E_t \Delta FS_{t+n} + \lim_{n\to\infty} \beta^n E_t B_{t+n}$$

[4] See Bohn (2007) for formal proofs.

arbitrary order, and a co-integration between these relevant variables is not a necessary condition for the satisfaction of the IBC.

Fiscal policy sustainability, however, depends on various and transitory impacts which may make the attempts of the governments to satisfy an IBC condition inconvenient (difficult). In this regard, Barro's (1979, 1986) tax smoothing model suggests governments can benefit from budget deficits (surpluses) in the short-and-medium period to smooth the tax rate in the case of temporary increases (decreases) in public expenditure, thus reducing the taxation distortionary effects. Therefore, it is important to question if these reactions in government budget balances compensate the impact of these various, mostly transitory shocks, hence ensuring the debt to follow a sustainable path. Within this scope and regarding the weaknesses of previous research for capturing fiscal sustainability, Bohn (1998, 2007) proposes an alternative test (fiscal reaction function) which has become increasingly popular among others.

Bohn's sustainability measure is based on a time-series estimation of the budget surplus on the public debt and other relevant variables. It measures to which extent policymakers respond to rising debt level that makes the primary balance better. The country's fiscal policy is considered as sustainable when this response is positive and statistically significant. A positive and significant reaction implies that a government is taking measures either by reducing primary expenditures or increasing revenues to offset the changes in public debt. However, this reaction can be concealed by outliers, such as wartime spending and fluctuations. He suggests that if one corrects these outliers, fiscal reaction function (FRF) approach is more promising than previous research that is based on stationarity and co-integration tests for fiscal sustainability analysis. The analytical framework for FRF is as follows:

$$s_t = pd_t + \omega Z_t + \epsilon_t \tag{10}$$

Where s_t is the primary surplus in GDP percent, d_t is the public debt in GDP percent at the beginning of the period. Z_t includes other determinants for fiscal primary balance; ϵ_t is a mean zero error term captures unobservable determinants of the primary surplus.

It is also important to find consistent control variables (Z_t) concerning the omitted variables issue. For this reason, Bohn explicitly follows Barro's (1979) tax-smoothing model employing two control variables: temporary increases in public expenditure (GVAR) and temporary fluctuations in the business cycle (YVAR).

Bohn's (1998) estimation suggests that if the coefficient of debt (p) is positive and still significant then the fiscal policy in general (and debt) satisfies the intertemporal budget constraint under uncertainty. Furthermore, if the condition p>(r-g)/(1+r) holds, where r and g denote the interest rate and growth rate

as percentage of output respectively, it is concluded that debt is stationary, or fiscal policy reacts strong enough to make the debt ratio gradually falls back to its original level. Overall, the basic intuition is that the primary fiscal balance should improve adequately in response to rising debt-to-GDP ratio. On the other hand, the test is robust to deviations in interest rates, debt structure, and growth rates (Mauro et al., 2015).

Even though Bohn's approach has a meaningful theoretical basis, there are also limitations in terms of interpreting coefficients, which have been noted in the studies of Mendoza and Ostry (2008) and Mauro et al. (2015). Mendoza and Ostry (2008) highlight that a positive primary balance response to debt ($p>0$) in the specification is sufficient but not necessary for fiscal sustainability. For example, in the short run governments may appear inefficient to react to rising debt for many years but they may improve their fiscal balance over an infinite horizon. On the other hand, finding $p>0$ does not ensure that a government will pay its debt in the long run. Therefore, a significant and higher p is not necessarily an indication of better fiscal solvency or sustainable policies. While interpreting the findings, it is also crucial to take shortcomings of single equation specification approach into account, where independent variables may endogenously be determined with the fiscal surplus.

Additionally, Mauro et al. (2015) highlight some other limitations. First, for some combinations of the debt ratio and the coefficient p, primary surplus suggested by the estimated fiscal policy response function may be too high to be politically attainable or realistic. Second, the test generally assumes an increasing debt-to-GDP ratio which clearly does not correspond to the truth for many countries. Thus for these countries, a negative and insignificant Bohn's coefficient does not necessarily indicate the lack of fiscal sustainability, but this can be specified as over accumulation of assets (Mauro et al., 2015). Finally, Bohn's (1998) criterion is not applicable for short-horizon evaluation as the regression needs a long historical information. However, policymakers or global investors are generally more interested in whether fiscal policy response is adequate over a shorter period.

2.5. Prior Empirical Strategies and Studies

Although there is no unified practical way to measure sustainability, the fiscal reaction function approach is gaining broader acceptance after the seminal work of Bohn (1998, 2007, and 2008). However, the FRF literature is heterogeneous regarding the dependent variables, explanatory variables and applied empirical strategies. Considering dependent variable, aside from the using actual primary balance (Bohn, 1998; Abiad and Ostry, 2005; Mendoza and Ostry, 2008; Ghosh et al., 2013; Mauro et al., 2015), several studies also use general government cyclically adjusted fiscal balance with or without lag as the dependent variable (Galí and Perotti, 2003; IMF, 2003; Debrun et al., 2008; Ghosh et al., 2011).

In terms of the non-debt explanatory variables, some studies use straightforwardly Bohn's specification, i.e. GVAR and YVAR, while others incorporate a wide range of various transitory determinants such as the output gap, expenditure gap, inflation, fuel and non-fuel commodity price, and proxies for political stability, institutions, default or restructuring years (IMF, 2003; Abiad and Ostry, 2005; Mendoza and Ostry, 2008; Ghosh et al., 2013). The reason is either to account for the transitory and cyclical impact of factors, such as business cycle and commodity price cycles, or to obtain the direct impact of these explanatory variables on primary balance, such as inflation, political stability and interest payments.

The FRF literature also shows substantial variety regarding the estimation strategies in cross-country analysis. Besides fixed effect (FE) being the most widely used estimation strategy (Galí and Perotti, 2003, Abiad and Ostry, 2005, Mendoza and Ostry, 2008, Ghosh et al.,2011; Mauro et al., 2015), there are also other methods applied such as IV, GMM, GLS, FGLS, Pooled OLS, MG, PMG, LSDVC estimator and Bayesian time-varying coefficients technique (TVC) (IMF, 2003; Abiad and Ostry, 2005).

Bohn's (1998, 2008) own works analysed long-run historical time-series data for the US over the period 1916–95 and 1793–2003. He concluded that the conditional reaction of primary fiscal surplus to public debt is positive and significant which is presented as evidence for sustainable fiscal stance. Several works have examined Bohn's fiscal reaction function on the national level. While some studies suggest a negative or insignificant response coefficient especially for France, Spain and Portugal, other studies find that countries react strongly to increasing public debt ratio in line with Bohn (Berti et al., 2016).

Following, Bohn's approach has also been applied to cross-country analysis by several studies. Mendoza and Ostry (2008) analysed fiscal solvency in a panel of 22 industrial countries for the periods 1980–2005 and 34 emerging market countries for the period 1990–2005. They use FE estimator with robust standard errors and country AR (1) coefficients. Their findings indeed support Bohn's (1998) results of a robust positive relationship between primary fiscal surpluses and public debt for both economies. They also separate each country into subgroups of high- and low-debt-to-GDP level economies and conclude that the countries with moderate debt levels are fiscally solvent but the solvency criterion is not satisfied for those with debt-to-GDP ratios surpassing the sample mean and median of each country group[5]. Moreover, the coefficient in response to rising debt level has been found stronger in emerging countries than advanced economies. The reason is that the emerging markets need to react stronger to sustain fiscal solvency in a riskier financial environment. Therefore, a higher reaction coefficient does not mean a higher fiscal sustainability, but it indicates that emerging countries

[5] The response coefficient for those countries is not significantly different from zero.

can afford relatively smaller mean debt levels in equilibrium. However, as Mauro et al. (2015) point out that apart from their dataset being relatively small, their analysis assumes a constant slope coefficient across developing and advanced economies. This is an important limitation, as it is widely rejected in data.

Ghosh et al. (2013) apply their framework motivated by Bohn's FRF to a panel of 23 advanced countries over 1970-2007. They use fixed effect estimator with robust standard error and country AR (1) error term process. By using polynomial cubic functions, they explore that the marginal reaction of fiscal balance in response to increasing debt level is non-linear. At moderate debt levels, the marginal response of the fiscal balance to debt is positive and significant; however, it tends to decline when the debt level increases to beyond a certain very high level (around 90-100 percent of GDP) despite remaining positive. These findings are presented as fiscal fatigue.

Finally, Mauro et al. (2015) analyse the sources of variation, in how fiscal policy reacts to rising government debt level both over time and across countries for 55 countries in the sample over the past 200 years. For country specific analysis, they relax Bohn's fixed long-term fiscal policy response assumption by applying structural break tests and rolling windows to explore variation in ρ. The application of Bai and Perron (1998) structural break tests allow the break points to be endogenously determined by fiscal data rather than being set a priori. These tests can provide evidence that policy response to changes in debt-to-GDP ratio often differs over time within a given country. Additionally, they use rolling windows which facilitate to assess fiscal prudency/profligacy based on only the data for specific periods.

For the panel estimation, they attempt to uncover the economic fundamentals that drive changes in the fiscal response by relaxing the assumption of a basic linear relationship where only one parameter p completely gauges fiscal policy reaction and remains constant over time. They analyse whether the fiscal policy reacts to unpredictable deviations in the long-term growth and the real interest rates and they interact them with the debt ratio. As underlying determinants of variations in fiscal responses Mauro et al. (2015) distinguish several economic and political fundamentals. Using fixed-effect panel estimation method with robust standard errors (clustered by country), they find that a negative growth shock raises fiscal surplus only slightly, whereas increases in the cost of borrowing leads to more pronounced fiscal policy response for a given increase in public debt. Inflation is negatively and significantly related to the budget surplus. The paper claims that augmenting the regression with these additional economic determinants (unexpected changes in long-term, sovereign debt costs and inflation), the goodness-of-fit of the panel regressions increases considerably.

Overall, the existing literature generally finds evidence of significant and positive response coefficients in response to increases in public debt ratio in cross-country analysis. In a broad literature review, Checherita-Westphal and Zdarek (2015) indicate that the magnitude of the response coefficients ranges between 0.01 and 0.10. Moreover, fixed effect estimator is extensively applied in the analysis. However, most of previous studies do not take into account the possible weaknesses of using FE estimator especially when cross-sectional dependence and parameter heterogeneity are ignored. FE methods may not be applicable in the FRF panel studies. In the next chapter, we will take a closer look at the potential econometric issues in the available literature and present another possible methodology to estimate a FRF for this cross-country analysis.

3. ECONOMETRIC ISSUES[6]

The primary aspects of the estimation strategy in this study are to model the empirical process being common and different across economies. The meaning of 'common' in this case is rather related to global shocks (i.e. the oil crisis in the 1970s or the most recent financial crisis originated in the USA in 2007) or their worsening impacts. In the literature, this commonality is identified as 'cross-sectional dependence'.

As it concerns econometric methodology, potentially important objections, largely ignored in the previous studies in FRF literature, are the implicit assumptions of cross-sectional independence (although they attempt to account for it, they do not formally test) and parameter homogeneity across countries. However, there is a growing consensus that panel data models often inherit cross-sectional dependence and parameter heterogeneity.

Cross-sectional dependence may emerge for a number of reasons such as spill over or spatial effects; or due to unobserved common components that influence all countries but most likely to a different extent, and ultimately enter into the error term (Coakley et al., 2006; Hoyos and Sarafidis, 2006 Pesaran and Tosetti, 2010). In the presence of various global shocks, the most recent financial crisis originated in the USA, addressing dependence is important, as these shocks often lead to synchronized fiscal responses across countries in traditional panel econometric methods.

Panel data literature addresses cross-sectional dependence in several ways, among which characterising cross-sectional dependence through a factor stands out as the most robust approach. The first approach is the standard spatial econometric modelling. This approach can be used in the case that the components of the correlation such as distance and location are known; thus, the cross-section dependence can be explicitly accounted in the models. However, the practical implementation of this method may not be

[6]I closely follow Eberhardt and Teal, 2011; Eberhardt, 2012 and Eberhardt Presbitero, 2013 for the structure of empirical model presentations.

applicable in FRF cross country analysis since it is restricted by the time-invariance spatial correlation (i.e. location or distance does not change over time).

The second approach is the use of fixed effects and time dummies to control for time-variant and time-invariant correlation across countries (Eberdhardt and Teal, 2011). However, this method explicitly imposes that all countries are identically affected by the impact of the cross-section components (i.e. global shocks), which is clearly not true. It is also important to underline that inclusion of year dummies prevent each country from having its own slope coefficient[7] (Eberdhardt and Teal, 2011). Therefore, it cannot fully capture the influence of the cross-sectional dependence. Furthermore, if unobservable components that lead to interdependencies across cross-sections are correlated with the independent variables, fixed effects estimators will be biased and inconsistent (Chudik and Pesaran, 2015).
The third and more promising approach is the application of a multi-factor error structure. If unobserved common factors that cause cross sectional dependence are not accounted accordingly resulting parameter estimates will be biased. To illustrate, the general specification in FRF literature (10) with a simple factor model[8] can be used:

$$s_{it} = p_i d_{it} + u_{it} \qquad u_{it} = \gamma_i + \theta_i f_t + e_{it} \tag{11}$$

Where p_i is the country-specific long-run parameter of interest, u_{it} is the error term including an unobserved country-specific fixed effect γ_{it} such as location and climate, and an unobserved time-varying common factor f with country-specific factor loading θ_i such as institutions, intangible capital, global shocks which influence all countries possibly to a different degree. The later could also be productivity spill overs in neighbouring countries.

Modelling d_{it} as a linear function of unobserved common factors f_t and g_t with country-specific factor loadings (ϑ_i and n_i) gives us:

$$d_{it} = \vartheta_i f_t + n_i g_t + v_{it} \tag{12}$$

Where e_{it} and v_{it} are white noise error terms (normally distributed with mean zero and constant variance $N(0, \sigma^2)$). In this setup, the unobservable factor f_t drives d_{it} in equation (12), and it is also a determinant of the error term u_{it} in equation (11), leading to cross-section dependence in the observables and unobservables. Therefore, d_{it} and u_{it} will be correlated in equation (11) and if cross-

[7] For example, there is T observations for each country, the regression is $Y_{jt} = b_j X_{jt} + \sum_2^T D_t + u_{it}$ would involve an intercept, a covariate and T-1 year dummies. This will add up T+1 to the regressors; thus, there is no degree of freedom.
[8] For simplicity, we only keep d_{it} as a regressor and adopt only a single unobserved component f.

sectional dependence is not accounted for, the parameter of interest (p_i) may not be identified. To illustrate, we can solve the equation (12) for f_t and replace this into the equation (11).

$$f_t = \vartheta_i^{-1}(d_{it} - n_i g_t - v_{it}) \tag{13}$$

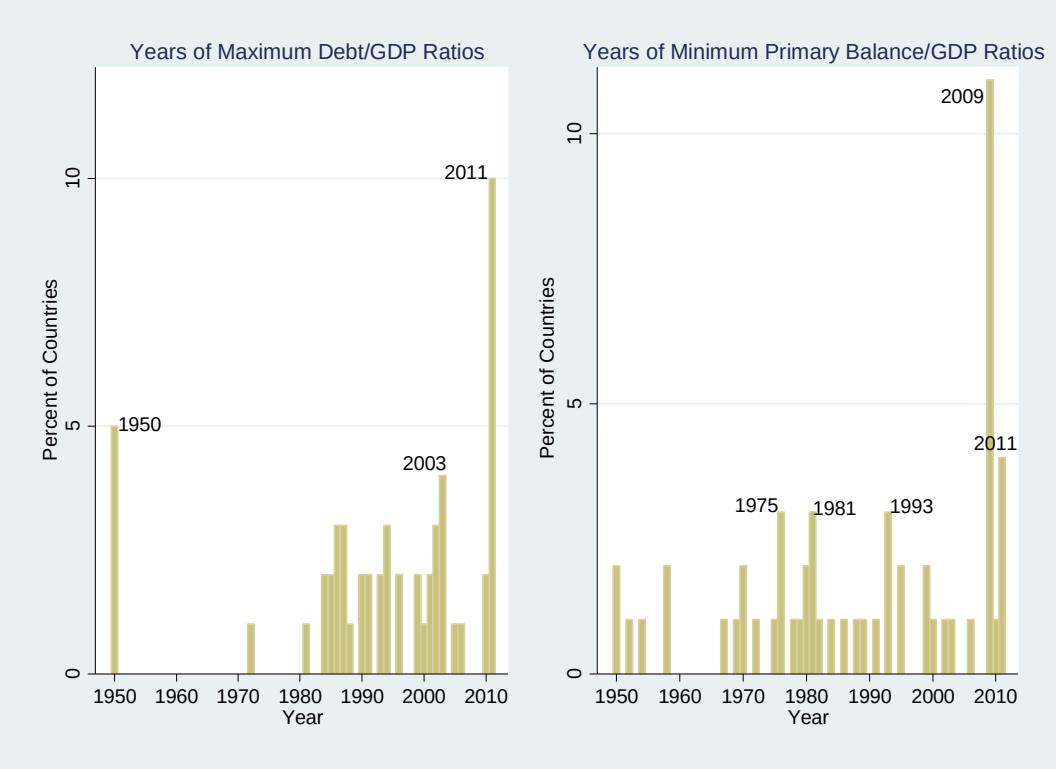
$$\begin{aligned} s_{it} &= \gamma_{it} + p_i d_{it} + \theta_i f_t + e_{it} \\ &= \gamma_{it} + p_i d_{it} + \theta_i \vartheta_i^{-1}(d_{it} - n_i g_t - v_{it}) + e_{it} \\ &= \gamma_{it} + \left(p_i + \theta_i \vartheta_i^{-1}\right) d_{it} - \theta_i \vartheta_i^{-1} n_i g_t - \theta_i \vartheta_i^{-1} v_{it} + e_{it} \end{aligned} \tag{14}$$

$$s_t = \gamma_{it} + B_{it} d_{it} + K_{it} \tag{15}$$

Where $B_{it} = \left(p_i + \theta_i \vartheta_i^{-1}\right)$ and $K_{it} = -\theta_i \vartheta_i^{-1} n_i g_t - \theta_i \vartheta_i^{-1} v_{it} + e_{it}$. From equation (14), it is apparent that standard panel estimators only yield a consistent estimator of $B_{it} = \left(p_i + \theta_i \vartheta_i^{-1}\right)$, or its average; we cannot obtain a consistent estimator of pure p_i (p_i is unidentified). Hence, the parameter of interest p_i cannot be identified without distinguishing it from the country-specific factor loadings such as θ_i, n_i, and the common components such as f_t and g_t.

Having motivated the concept of commonalities (cross-sectional dependence), we present some descriptive statistics with the purpose of seeking evidence for the potential cross-section dependence in debt/GDP ratio and primary balance across countries over the sample period 1950 to 2011. The data and sources are given in detail in the chapter (4). Figure 1 presents a histogram for each year over the sample in which the countries exhibit the peak debt ratio and minimum primary balances in GDP percentage in that year. Although the sample coverage differs across countries for this period, the figure shows those peak years for more than 34 percent of the countries coincide with only four years, namely, 1950, 2003 and 2011 for debt ratios. For more than 45 percent of the countries, the worst primary balances occurred only in a few years, namely, 1975, 1981, 2009 and 2011. Having said that the data covers over sixty years, figure 1 suggests evidence for cross-sectional dependence across countries that the debt ratio peaks and minimum primary balance to GDP ratio years are crowded around a smaller number of years.

Figure 1. Peak/Minimum Years

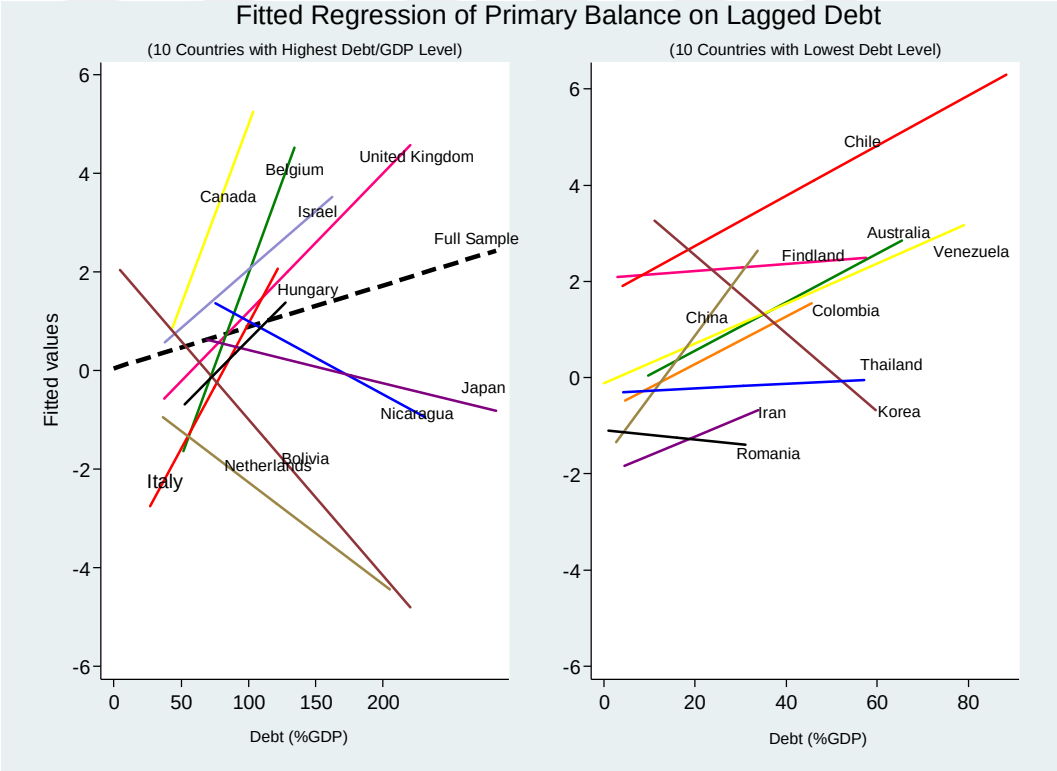


Empirical exercises in FRF literature also heavily rely on estimation methods (i.e., fixed effect models) which usually allow countries to have only different country-specific fixed effects (varying intercepts), but impose an implicit assumption of slope homogeneity across units. This is the simplest but very restrictive attempt to account for heterogeneity in a panel. However, when one studies complex economic behaviours such as countries, assuming slope parameter homogeneity across sections appears to be inappropriate (Durlauf et al., 2001, Brock and Durlauf, 2001 and Eberhardt and Teal, 2011). The potential problems that slope heterogeneity causes in standard panel data estimations (fixed effects) are highlighted in the study of Juhl and Logovskyy (2013). They mainly emphasise three major concerns on slope heterogeneity when using FEE. First, in the presence of slope heterogeneity across units, the standard fixed effect estimator will be inconsistent. Asymptotically, fixed effects estimator corresponds to the weighted average of slopes, which does not necessarily reflect the average population slope[9]. Thus, in the case of slope heterogeneity, substantial discrepancies between these two averages can emerge resulting in quantities that are different from parameters of interest. Second, even if these two means collapse to each other asymptotically, such an average quantity does not easily convey a useful and an immediate conclusion about estimated parameters. For instance, is it correct to conclude that an estimated average zero slope coefficient mean that the variable has no effect? In fact, the effect can be positive for some countries and negative for others. If it is the case, FEE will give insignificant

[9] Weights depend on variability of dependent variables in each cross-sectional unit.

coefficients while they are significant for some cross-sectional units. Third, similar to the case when we ignore individual intercepts in fixed effects, where omitted variables enter into the error terms, disregarding heterogeneous coefficients will also distort standard errors resulting in incorrect inference. Bearing in mind these concerns, in figure (2) we introduce a graphical illustration plotting linear regression lines for primary balance against debt-to-GDP ratio for countries with highest (lowest) debt-GDP ratios over the period 1950 to 2011. The graphs give a strong suggestive evidence for substantial heterogeneous fiscal response coefficients across countries. While some countries react strongly to increases in debt ratio, the responses of the primary balance become insignificant or weaker in other countries. Overall, In the light of the above discussion, it is, therefore, important to allow each country to have its own debt coefficient in the regression.

Figure 2. Linear Regression Lines



4. DATA, MODEL AND ECONOMETRIC STRATEGY

4.1. Data

This study is based on an unbalanced panel of 55 countries consisting of 31 non-advanced and 24 advanced economies over 1950-2011. It is composed of largely OECD countries. Non-OECD countries include Argentina, Brazil, Colombia, India, Peru and Russia. The dataset mainly covers: gross public debt, government expenditure and primary balance as shares of GDP percentages, real GDP, real growth rates, interest rates and inflation. The coverage of country-specific periods in non-advanced economies varies relatively depending upon the availability of data. Table 1 shows the list of all countries and corresponding periods. While the country coverages start from 1950s for most of the advanced countries, these are more recent and varying for most of the non-advanced countries.

Table 1. Country Coverage of Main Variables (Debt, Primary Balance and macro variables)

Country	Observation Number	Coverage	Country	Observation Number	Coverage
Argentina	62	1950-2011	Italy	62	1950-2011
Australia	62	1950-2011	Japan	62	1950-2011
Austria	62	1950-2011	Mexico	62	1950-2011
Belgium	62	1950-2011	Netherlands	62	1950-2011
Bolivia	57	1955-2011	New Zealand	62	1950-2011
Brazil	39	1965–1986, 1995–2011	Nicaragua	20	1988–2001, 2005–2011
Bulgaria	17	1995-2011	Norway	62	1950-2011
Canada	62	1950-2011	Pakistan	58	1951–1967, 1971–2011
Chile	45	1977-2011	Panama	52	1956–1977, 1982–2011
China	25	1987-2011	Paraguay	45	1967-2011
Colombia	62	1950-2011	Peru	42	1970-2011
Costa Rica	56	1956-2011	Philippines	58	1954-2011

Denmark	58	1954-2011	Poland	38	1984-2011
Dominican Republic	32	1980-2011	Portugal	62	1950-2011
Finland	62	1950-2011	Romania	32	1980-2011
France	62	1950-2011	Russia	14	1998-2011
Germany	62	1950-2011	South Africa	59	1950-1976, 1980-2011
Ghana	50	1962-2011	South Korea	54	1958-2011
Greece	62	1950-2011	Spain	62	1950-2011
Haiti	12	2000-2011	Sweden	62	1950-2011
Honduras	60	1952-2011	Switzerland	62	1950-2011
Hungary	28	1984-2011	Thailand	55	1957-2011
Iceland	62	1950-2011	Turkey	42	1970-2011
India	62	1950-2011	United-Kingdom	62	1950-2011
Indonesia	40	1972-2011	United States	62	1950-2011
Iran	35	1962–1965, 1980–2011	Uruguay	40	1972-2011
Ireland	62	1950-2011	Venezuela	54	1950-1979, 1988-2011
Israel	49	1954–1970, 1980–2011			

The data providing primary balance, government expenditure, primary balance and gross public debt expressed as shares of GDP, real long-term interest rates on public debt, and output growth rates used in the study come from the IMF's historical public debt database compiled by Mauro et al. (2015).

For almost all countries in the sample[10], real GDP data series are drawn from The World Economic Outlook (WEO) database of IMF. Therefore, real expenditure series are computed by using real GDP series and expenditure-to-nominal GDP series from the dataset of Mauro et al. (2015). Following Mauro et al. (2015), interpolated expenditure values are used for real expenditure data gaps of three years or less. Besides, there are unnatural changes by cause of shifts in coverage of the levels of government expenditures which are shown by dummy variables in the dataset of Mauro et al. (2015). In this case, all the values for general government real expenditure were scaled by the ratio of central government expenditure over general government expenditure in the year of the sector coverage shift.

[10] For Iceland the real GDP series is taken from Federal Reserve Bank of St. Louis.

Inflation series based on the GDP deflator are sourced from the WEO. While the coverage starts in the 1960s for most of the developed countries, it is more varying and recent for the developing countries.

Figure 3 represents the evolution of the primary balance (top panel) and gross public debt (bottom panel) as shares of GDP percent over the period 1950 to 2011 for all countries. The bars represent the 95% confidence interval. It is apparent from the graphs that while the gross public debt ratio continuously decreases until the early 1970s, relatively more stable and positive average fiscal balances are observed. After the 1970s, average debt ratios generally keep increasing with some fluctuations, and the average fiscal balances exhibit slight further improvements. Particularly, there are substantial improvements in the average primary balances in the 1990s corresponding to financial consolidation efforts such as fiscal adjustment plan in Canada, the Maastricht euro-entry process in several Eurozone. The structural primary balance deteriorations take place in the 1970s, early 1980s, 1990s, and 2000s. The most recent financial crisis in 2007 caused the most highly influential deterioration of the primary balances in 2008 and 2009 over the period 1950 to 2011.

Figure 3. Evolution of Debt

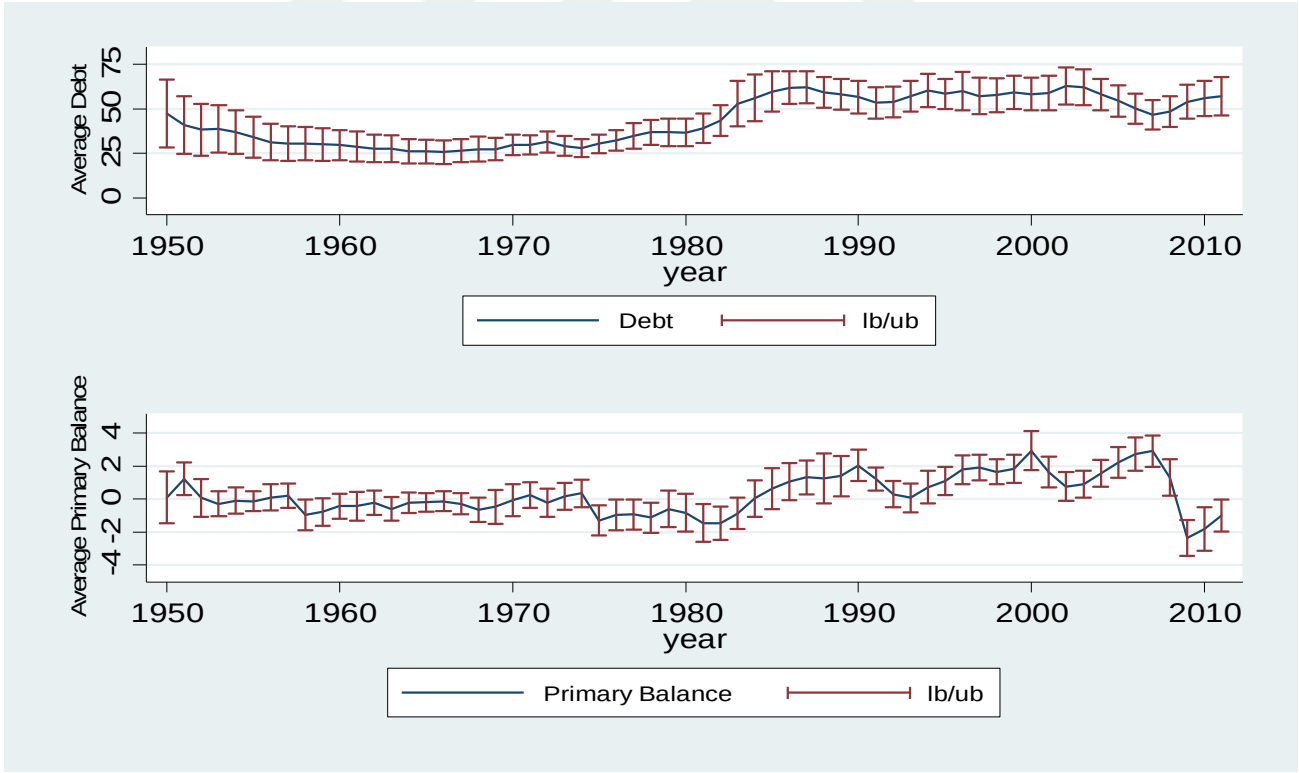


Table 2 reports summary statistics for the full sample period (1950−2011) including the primary balance and total government public debt. The data consists of approximately 2750 observations for each of primary balance and debt for all countries.

Table 2. Summary Statistics

	All Countries									
	Mean	Std. Dev.	Min	25%	Median	75%	95%	99%	Max	Obs.
Primary Balance	0.3 -	3.5	-28.2	-1.5	0.3	2.2	5.9	9.6	20.6	2803
Debt	45.1	34.3	0.1	19.7	37.8	60.6	106.7	169.9	289.6	2983
	Non-Advanced Countries									
	Mean	Std. Dev.	Min	25%	Median	75%	95%	99%	Max	Obs.
Primary Balance	0.0	3.4	-22.6	-1.9	-0.1	1.7	5.7	9.0	18.2	1336
Debt	41.0	34.2	0.1	17.6	33.1	54.7	99.8	188.8	289.6	1525
	Advanced Countries									
	Mean	Std. Dev.	Min	25%	Median	75%	95%	99%	Max	Obs.
Primary Balance	0.7	3.5	-28.2	-1.1	0.6	2.5	6.0	10.1	20.6	1467
Debt	49.4	34.0	1.6	22.9	44.3	64.6	113.2	166.2	284.0	1458

Source: Mauro et al. (2013) Working Paper

For the fiscal balance, the average is around 0.3 in percent of GDP and it is 45.6 in percent of GDP for the total debt ratio. Primary balances are more than 9.1 percent of GDP in the 99th percentile of the distribution, which mostly includes the countries with rather large government assets or commodity producing countries. Primary balances in the 95th percentile are in excess of 5.8 in percent of GDP corresponding to those countries with smaller government assets. In comparison to the non-advanced economies, advanced countries display greater debt-to-GDP ratios and primary balances.
Default years are taken from Reinhart and Rogoff (2009) and are excluded from the analysis. Especially Latin American countries experienced defaults largely in the 1970s and 1980s.

4.2. Model Specification

Our theoretical model and empirical specification closely follow the recent work of Mauro et al. (2013) which is motivated by Bohn's (1998) argument. Given the discussion in Chapter 2 and 3, we represent an equation introducing the primary determinants impacting the fiscal outcome in the period 1950-2011. Differently from Mauro et al., what we empirically add to their model is the use of more flexible and recent empirical techniques that accounts for cross-sectional dependence and parameter heterogeneity in the residuals in order to identify the parameter of interest. This section only serves to explain the theoretical model specification; our augmentation will be detailed in the next section.

As documented in the section (2), Mauro et al. have shown that the degree of the fiscal reaction coefficient to variation in debt-to-GDP shows substantial differences across countries and over time, and they explore the drivers underlying such differences. To do this, they relax Bohn's (1998) assumption that the single coefficient estimate *p* completely captures fiscal policy reaction and remain constant over time after controlling for cyclical components for output and expenditure; and they allow policy response to depend on unpredictable deviations in the long-term growth and the real interest rates. The intuition behind is that if politicians do not realize the permanent slowdown in economic growth, they may fail to improve the primary fiscal balance for a given rise in the debt ratio. Besides, politicians will improve the primary fiscal balance in case of a rise in the borrowing cost (real interest rate) for a given increase in the debt ratio.

In particular, we first estimate Bohn's fiscal reaction function:

$$pb_{it} = p * d_{it} + \omega_{it} * Z_{it} + u_{it} \tag{16}$$

Allowing the response to increases in debt to depend on unpredicted declines (inclines) in long-run growth rates (interest rates), the equation (16) takes the form:

$$\begin{aligned} pb_{it} &= \underbrace{[p(1 + c_1 g_{it} + c_2 r_{it})]}_{varying\ response\ to\ debt} d_{it} + \omega_{it} Z_{it} + u_{it} \\ &= pd_{it} + pc_1 g_{it} d_{it} + pc_2 r_{it} d_{it} + \omega_{it} Z_{it} + u_{it} \end{aligned} \tag{17}$$

When we include also the direct impacts of the growth surprises and interest rates, we obtain:

$$pb_{it} = pd_{it} + \underbrace{pc_1}_{a1} g_{it} d_{it} + \underbrace{pc_2}_{a2} r_{it} d_{it} + \underbrace{c_3}_{a3} g_{it} + \underbrace{c_4}_{a4} r_{it} + \omega_{it} Z_{it} + u_{it}$$

$$pb_{it} = pd_{it} + a_1 g_{it} d_{it} + a_2 r_{it} d_{it} + a_3 g_{it} + a_4 r_{it} + \omega_{it} Z_{it} + u_{it} \tag{18}$$

Where pb_{it} is the primary balance in GDP percent, d_{it} is the debt-to-GDP, g_{it} is the growth surprise (i.e. the unpredicted deviation in the long-term growth), r_{it} is the real long-term interest rate, and Z_{it} includes other observable determinants of the primary fiscal balance such as output gap and expenditure gap.

The first interaction term is the combination of debt ratio and the unexpected deviation in the long-term growth rate. The second interaction term combines debt-to-GDP ratio with the real long-term interest rate (borrowing cost) on newly issued debt. In this regard, if a_1>0 and a_2>0, it is interpreted as that a

positive (negative) deviation in real interest rate (growth shock) marginally improves (weakens) the fiscal policy response to increases in debt ratio.

To calculate the growth surprise variable, we first make five-step ahead auto-regressive dynamic forecasts of GDP growth rates for each country-year, rolling the sample recursively, with the starting window size of 20. Then based on these forecasts we project real GDP variables for next 5 years for each country-year. Next, for each-country year, we generate GDP series from actual GDP series up to the known date plus 5-year forecasts. This reflects policy makers' beliefs under partial information (based on historic information up to the realized date) about future levels of output for next 5 years. Then we HP trend later series using smoothing parameter "6.25" which is standard for annual data. Finally, we HP filter actual real GDP series over the full sample. The growth surprise variable is the difference between trend growth rates under partial information and full sample information. In effect, it measures the deviation of potential long run growth (based on actual, full sample information) from long run growth based on beliefs of policy makers (for each country-year). The long-run borrowing cost r_{it} for each country is the ten-year domestic currency sovereign borrowing rate minus GDP deflator inflation.

The output and expenditure gaps (i.e. the cyclical components) are computed by de-trending those series utilizing the Hodrick-Prescott filter with the smoothing parameter value of 100. The Hodrick-Prescott filter was applied to the separate periods for countries with 25 years of available expenditure data separated by at least a four-year gap.

4.3. Estimation Strategy

We employ Pesaran's (2006) Common Correlated Effects Mean Group estimator (CCEMG). The purpose of utilising this model setup that diverges from the standard panel estimators is two-fold. First, it accounts for the potential cross-country parameter heterogeneity in the impact of observables and un-observables. Second, this setup addresses the possible cross-section components of the residuals, which ultimately causes endogeneity problem. In this method, the unobserved component f can be controlled for if the sample size and time dimension are moderate or sufficiently large. To be more specific, we take the cross-section average of the equations (11), (12) and (13) from section (3) and replicate in this section:

$$pb_{it} = \gamma_{it} + p_i d_{it} + \theta_i f_t + u_{it} \tag{19}$$

$$\overline{pb_t} = \bar{\gamma} + \bar{p}\overline{d_t} + \bar{\theta} f_t \qquad \text{given that } \overline{u_t} = 0 \text{ as } N \to \infty \tag{20}$$

$$f_t = \bar{\theta}^{-1}\left(pb_t - \bar{\gamma} - \bar{p}\overline{d_t}\right) \tag{21}$$

Where cross-section means at time t are specified as $\overline{pb_t} = N^{-1}\sum_{i=1}^{N} pb_{it}$ and $\overline{d_t} = N^{-1}\sum_{i=1}^{N} d_{it}$.[11] As N becomes large, equation (19) can be solved for the cross-sectional dependence through a linear combination of the cross-sectional averages of dependent and independent variables. Replacing $f_t$in equation (19) yields:

$$pb_{it} = \gamma_{it} + p_i d_{it} + \theta_i\bar{\theta}^{-1}\left(pb_t - \bar{\gamma} - \bar{p}\overline{d_t}\right) + u_{it} \tag{22}$$

$$pb_{it} = p_i d_{it} + \pi_{1i}\overline{pb_t} + \pi_{2i}\overline{d_t} + \pi_{3i} + u_{it} \tag{23}$$

The CCEMG estimation method generates an unbiased estimation of the heterogeneous long-run parameter of interest ($p_i \neq p$) where the cross-section averages for the dependent ($\overline{pb_t}$) and independent ($\overline{d_t}$) variables (for the entire panel) act as placeholders in the country-specific regression equation.

However, as pointed out Eberdhardt (2012), the focus of the Pesaran (2006) is to achieve consistent estimation of the model parameters regarding the observables. Empirically, estimated coefficients for the cross-section-averaged variables (pb_t and $\overline{d_t}$) and their average estimates (π_{1i} and π_{2i}) cannot be interpreted in a meaningful way; they are employed for only to correct the biasing impact arisen from the unobservable common factor.

In the analysis of fiscal and macroeconomic data that are very likely to have strong inter-dependence, assuming cross-sectional independence appears to be inappropriate. Therefore, before and/or after conducting empirical analysis, it is also important to test the presence of cross-sectional dependence. There are three present tests for examining cross-sectional dependence in the present panel econometrics literature, i.e.; Friedman (1937), Frees (1995) and Pesaran's (2004) cross-sectional dependence (CD) tests. In this paper, the null of the cross-sectional independence was tested by employing the Pesaran's (2004) CD test for unbalanced panels, which uses the correlation coefficients obtained from the residuals of individual regressions. Therefore, the hypothesis of interest is:

$$H_0: \mathrm{p}_{ij} = \mathrm{p}_{ji} = corr\left(u_{it}, u_{jt}\right) = 0 \; for \; i \neq j \tag{24}$$

[11] . We also construct cross-section averages for each in the specifications with multiple covariates: $\overline{gd_t} = N^{-1}\sum_{i=1}^{N} g_{it}d_{it}$, $\overline{rd_t} = N^{-1}\sum_{i=1}^{N} r_{it}d_{it}$, $\overline{g_t} = N^{-1}\sum_{i=1}^{N} g_{it}$, $\overline{r_t} = N^{-1}\sum_{i=1}^{N} r_{it}$ and $\overline{Z_t} = N^{-1}\sum_{i=1}^{N} Z_{it}$.

Where p is specified as the residuals' pair-wise correlation:

$$\hat{\rho}_{ij} = \frac{\sum_{t\in T_i\cap T_j}(\hat{u}_{it}-\bar{\hat{u}}_i)(\hat{u}_{jt}-\bar{\hat{u}}_j)}{\sqrt{\left[\sum_{t\in T_i\cap T_j}(\hat{u}_{it}-\bar{\hat{u}}_i)^2\right]}\sqrt{\left[\sum_{t\in T_i\cap T_j}(\hat{u}_{jt}-\bar{\hat{u}}_j)^2\right]}} \quad (25)$$

Therefore, the CD statistics for unbalanced panels is set by:

$$CD = \left[\frac{2}{N(N-1)}\right]^{\frac{1}{2}}\left(\sum_{i=1}^{N-1}\sum_{j=i+1}^{N} T_{ij}^{1/2}\hat{\rho}_{ij}\right) \quad (26)$$

Where $T_{ij} =\neq (T_i \cap T_i)$.

5. RESULTS

This chapter presents results of the panel estimates of equation (23). The panel regressions are divided into three groups, namely, full country sample, non-advanced countries and advanced countries.
We first implement the fixed effect panel estimations with robust standard errors clustered by country groups. Given the issue of identification problem regarding cross-sectional dependence, which may arise from global shocks, spillovers or correlation in the residuals, the Pesaran (2004) CD test is conducted after estimates. The CD tests for all country groups are reported in the Appendix in Table 1A.1 to Table 1A.3. The Pesaran (2004) test clearly rejects the null hypothesis of cross-sectional independence in all the FE models presented in tables (CD test statistics are above 20, and all respective p values are 0).

5.1. Results for the Sample of All Countries

Following, we applied CCEMG estimation as shown in equation (23) using the model specification of equation (18), which addresses the potential parameter heterogeneity by allowing debt response coefficients to differ across countries and commonalities by controlling cross-section averages. Table 3 provides the results of four distinct model specifications that are labelled R1 to R4, respectively. All specifications in the table include output gap and expenditure gap to control the cyclical components and transitory impacts of government expenditure. Each specification is separated into two columns that are A and B. While A refers to the specification without trend, B represents the models with a group-specific linear trend. The primary focus in R1 is only the parameter of interest p (fiscal response coefficient) in line with Bohn's (1998) specification. R2 augments R1 with the interaction term of lagged debt and long-term growth surprise. R3 includes the direct impact of growth surprise on primary balance. Finally, R4 adds the GDP-deflator inflation rates.

Table 3. CCEMG estimator: Full Country Sample

[1]	R1		R2		R3		R4	
	A	B	A	B	A	B	A	B
[2]	.018 ***	.026 ***	.020 **	.039 ***	.026 ***	.042 ***	.026 **	.035 ***
[3]			.18 ***	.16 **	.22	.33 *	.16	.37 ***
[4]					-7.6	-14.2 *	-6.8	-11.8 *
[5]								0.003
[6]	.14 ***	.15 ***	.14 ***	.17 ***	.17 ***	.18 ***	.19 ***	.20 ***
[7]	-.14 ***	-.14 ***	-.19 ***	-.19 ***	-.20 ***	-.20 ***	-.21 ***	-.19 ***
[8]	-	included	-	included	-	included		Included
[9]	2692	2692	2032	2032	2032	2032	2005	2005
[CD]	**[-1.2]**	**[-1.29]**	**[-0.5]**	**[-0.30]**	**[-0.60]**	**[-0.01]**	**[0.49]**	**[-0.71]**
[p]	**[0.20]**	**[0.19]**	**[0.61]**	**[0.76]**	**[0.55]**	**[0.99]**	**[0.62]**	**[0.48]**

[1]: Primary Balance: pb_{it}

[2]: Lagged Debt: $p_i d_{it}$

[3]: Debt*Growth Surprise: $a_1 g_{it} d_{it}$

[4]: Growth Surprise: $a_2 g_{it}$

[5]: Inflation

[6]: Output Gap

[7]: Expenditure Gap

[8]: Trend

[9]: Number of observation

[CD]: CD-Statistics

[p]: p value

CCEMG estimator with standard robust errors: $pb_{it} = p_i d_{it} + a_1 g_{it} d_{it} + a_2 g_{it} + \omega_{it} Z_{it} + \pi_{3i} + u_{it}$. We do not report the coefficients for cross-section averages for space. Asterisks show significance at the 10, 5, and 1 percent levels, respectively.

Relying on residual diagnostic, we can observe that CD-statistics are drastically reduced for the CCEMG models (varying from -1.33 to -0.20) comparative to the FE models. Furthermore, while the null hypothesis of independence is still rejected in some specifications without trends, cross-sectional dependence is fully accounted with the inclusion of country-specific linear trends (R2, R3 and R4). Including trends allows taking into account any unobserved common factors that smoothly changes over time. Notably, the inclusion of trend reduces cross-sectional dependence further, and alters the significances and the magnitudes of the other results in almost all the specification aside from the first specification (R1). Hence, we consider the specification without trend R1A to interpret Bohn's coefficient and the ones with trend to evaluate potential variation in the fiscal policy response function to growth surprises and interest rate.

When considering R1A, the regression yields a positive debt coefficient (p) suggests that for a given increase of 100 basis point in the debt ratio, governments react by improving the primary balance-to-GDP ratio of 1.8 basis points. This fiscal response is statistically significant at the 99 percent confidence level. Comparing this finding with studies applied FE estimator, we find a lower Bohn reaction function once we apply CCEMG. Bohn's (1998) own estimates of debt coefficients vary between 0.028 to 0.054 in the full period and five subsample periods. Mendoza and Ostry (2008) finds this coefficient varying between 0.035 to 0.041.

The overall fiscal balance reaction to a marginal government debt increase is specified as ($p_i d_{it} + a_1 g_{it} d_{it}$) (eq. (18)). The estimations of a_1 are positive and significant at the .01 level for the specifications in R2A, R2B and R4B. That is to say, an unexpected deviation in the long-term growth rate marginally weakens the policy response of the primary fiscal balance to increases in debt-GDP ratio. The direct impact of long-term growth surprises specified as a_2 is independent of the debt-to-GDP ratio and it also enters significantly and positively in the specification R4B.

To illustrate the size impact of growth surprises on primary balance, based on the regression in R4B for a government facing a 100 percent debt level, a 10 percent increase in debt-GDP ratio leads to an increase in primary balance by 0.3 percent in the absence of long-term growth shocks. In case that real long-term growth rate unexpectedly decreases by 1 percent, a_1=0.32 (the interaction term) suggests that the response of the primary balance to the corresponding 10 percent of an increase in debt ratio declines by 3.2 percent (i.e. 100*0.32*-.01). The direct impact of a growth shock (a_2) indicates an immediate upward correction in the fiscal balance by 8.7 percent (-.01*-8.7). Thus, the overall policy response for a given increase in debt is 9 percent in the presence of a negative long-term growth shock.

The overall relationship between growth surprise variables and the primary surplus is given by $a_1 g_{it} d_{it} + a_2 g_{it}$. When it is the case: $a_1 g_{it} d_{it} + a_2 g_{it}$>0, the effect of the growth surprise will change the sign for countries with debt level $\hat{d}'=a_2/a_1$. For the full sample of advanced and non-advanced countries this threshold ˆd=0.27. In this manner, for the economies with debt level excess of 27 percent of GDP (the majority of the countries), negative deviations from the long-run growth path have given rise to greater fiscal insolvency between the period 1950 and 2011. Mauro et al. (2015) find this threshold level as 9 percent of GDP in their corresponding regression and 7 percent in the full specification for the same period. In comparison to our findings, their estimates suggest a much lower threshold debt level.

The results also show that the coefficient of inflation (R4) is not statistically different from zero when using the full sample. Nevertheless, including GDP-deflator inflation as an explanatory variable does not change the significance of the other variables.

Adding the interaction term of debt with inflation to the regression models was statistically insignificant; therefore, we have not reported the results in this exercise. Following Mauro et al. (2015) and Reinhart and Sbrancia (2015), the insignificance of the interaction term of debt and inflation can be interpreted in the following way: In the late 60s and 70s factors such as the inter-country capital mobility and regulatory constraints led to financial repression. Whereas, today concepts like open borders and markets and financial deregulation melted down the erosion of the true value of debt harder. However, the debt of the government from the 80s up to today became higher than in the 70s. The meeting point of high debts and low ability to back a financial repression may be the reason of the insignificance of the interaction term between debt and inflation.

We also differently than Mauro et al. have not included the interaction term between debt and interest rate in regressions for the all country sample due to the non-availability of the data. The data for the interest rates are mainly available for advanced economies, while it is available for only a few non-advanced countries. Thus, interpretation of the interaction term of debt with interest rates for the full sample can conceal relevant differences in the behaviour of fiscal and macroeconomic variables across countries in the two groups (advanced and non-advanced economies).

5.2. Results for the Sample of Non-Advanced Countries

Table 4 presents estimation results and CD-statistics using the sample of non-advanced economies in a similar set of model regressions as applied to the full country sample. We have reported only the models with trends. Focusing on the residual diagnostic, CCEMG estimator accounts for cross-sectional dependence in each model. Comparing to the FE estimator (presented in the appendix table 2A), CD-statistics are now much smaller (around -0.49 to -1.50) and all the respective p values are excess of 10 percent level suggesting enough evidence for failure to reject the null hypothesis of cross-sectional independence.

Table 4. CCEMG estimator: Developing Countries

Primary Balance to GDP	R1	R2	R3	R4
1950-2011				
Lagged Debt	.026	.53	.058	.024
	**	***	***	
Debt*Change in HP Growth		.26	.44	.43
		***	**	***
Change in HP Growth			-10.1	-6.58
Inflation				.009
				*
Output Gap	.12	.08	.08	.13
	***	**	*	***
Expenditure Gap	-.10	-.11	-.10	-.11
	***	***	***	***
Number of Observations	1249	1016	1016	1013
CD-Statistics	**-1.13**	**-1.16**	**-1.50**	**-0.49**
[p] value	**[0.25]**	**[0.24]**	**[0.13]**	**[0.62]**

CCEMG estimator with standard robust errors: $pb_{it} = p_i d_{it} + a_{1i} g_{it} d_{it} + a_{2i} g_{it} + \omega_{it} Z_{it} + \pi_{3i} + u_{it}$. We do not report the coefficients of cross-section averages for space. Asterisks show significance at the 10, 5, and 1 percent levels, respectively.

Turning on the estimation results, we can observe that the results for the non-advance countries have similar characteristics to those of the full sample country group. The point estimates of *p* ranges from 0.024 to 0.058 (slightly larger comparing to full sample coefficients) consistent with the IBC. The first

model, R1, implies that for a given increase of 10 basis point in the debt ratio is associated with a 0.26 percent increase in annual primary balance in GDP percent. This fiscal response is statistically significant at the 95 percent confidence level.

The point estimates of a_1 enter positively and significantly in all the regression models. Therefore, the interaction terms of debt with growth surprises have a significant association with the size of the policy response to increasing debt level in developing countries. To exemplify the size effect, consider R3 for a hypothetical country with a 100 percent debt-GDP ratio: in the case that the country negatively deviates 1 percent from its long-term growth path, the interaction term a_1 implies that the fiscal surplus declines by 44 percent of GDP (100*0.44*-.01). The direct impact of the growth surprise; however, make an upward correction in the primary balance by 10 percent (-10.1*-.01) (insignificantly). Therefore, the overall primary balance response to a marginal 10 percent increase in debt-GDP ratio expressed as in equation (18) ($p_i d_{it} + a_1 g_{it} d_{it} + a_2 g_{it}$) equals to 24 percent of GDP resulting from 0.58*100*0.1+100*0.44*-.01+-.01*-10.1. Accordingly, focusing on the overall link between the primary balance and growth surprise, given by $a_1 g_{it} d_{it} + a_2 g_{it}$, the estimation exhibits a smaller threshold debt-to-GDP ratio 23 percent ($\hat{d}$'=a_2/a_1).

5.3. Results for the Sample of Advanced Countries

Table 5 provides the results and CD-statistics for seven distinct model specifications that are labelled R1 to R7 using the sample of advanced countries. All model specifications in the table take account of output gap and expenditure gap to control the cyclical impacts and transitory components of government expenditure and country-specific trend.

Table 5. CCEMG Estimator: Advanced Countries

[1]	R1	R2	R3	R4	R5	R6	R7
[2]	.015	.014	.038**	.027*	.016	.058***	.037
[3]	-	-.11	0.43*	-	-	.29	.32
[4]	-	-	-27.1**	-	-	-21.2*	-26.4*
[5]	-	-	-	-.13	.18	.15	.56*
[6]	-	-	-		-.25*	-.14	-.37*
[7]	-	-	-	-	-	-	-.06*
[8]	.21***	.31***	.30***	.27***	.27***	.29***	.34***
[9]	-.19***	-.35***	-.36***	-.23***	-.24***	-.37***	-.40***
[10]	1443	1016	1016	1178	1178	904	880
[CD]	-3.05	-2.68	-2.04	-2.97	-2.50	-2.67	[-1.83]
[p]	[0.002]	[0.007]	0.041	[0.003]	[0.012]	[0.008]	[0.06]

[1]: Primary Balance
[2]: Lagged Debt
[3]: Debt*HP Growth
[4]: HP Growth surprise
[5]: Debt* Real LT rate
[6]: Real LT rate
[7]: Inflation
[8]: Output Gap
[9]: Expenditure Gap
[10]: Number of observation
[CD]: Cross Sectional Independence, Pesaran test

CCEMG estimator: $pb_{it} = p_i d_{it} + a_{1i} g_{it} d_{it} + a_{2i} g_{it} + a_{3i} r_{it} d_{it} + a_{4i} r_{it} + \omega_{it} Z_{it} + \pi_{3i} + u_{it}$. We do not report the coefficients of cross-section averages for space. Asterisks show significance at the 10, 5, and 1 percent levels, respectively.

R1 allows only the fiscal response coefficient p in line with Bohn's (1998) specification. R2 shows results adding the interaction term between debt level and growth surprise. R3 augments R2 with the direct impact of growth surprise beside its interaction term with debt level. R4 allows the fiscal response coefficient and its interaction with the interest rate. R5 adds the direct effect of real long-term interest rate to R4. R6 augments R5 with the interaction term between debt and growth surprise and the direct impact of growth surprise. Finally, R7 includes the direct effect of inflation on primary balance.
We can observe the corresponding model regressions using FE estimator suffer from substantial cross-section dependence (see appendix table 1A. 2.). Although it is drastically reduced in the CCEMG models, we still reject the null hypothesis of no cross-section dependence (CD-statistics are around -3.21 to -2.21). One potential explanation could be as follows: The data for developed countries mainly consist of European countries. The dependency in the data generation process to a certain extent could be arising from the strong interconnectedness of these countries through trade, institutions, politics or the common settings for financial regulation at the EU level.

The policy response to a marginal government debt increase is specified as ($p_i d_{it} + a_{1i} g_{it} d_{it}$) (eq. (18)). The estimation of a_2 is found to be positive and significant at the .10 level in model R3. It implies that the interaction term of debt with growth surprises have a significant association with the size of the policy response to increasing debt level in advanced countries. Subsequently, focusing on the overall link between the primary balance and growth surprise, given by $a_{1i} g_{it} d_{it} + a_{2i} g_{it}$, the estimation yields a higher threshold debt-to-GDP ratio of 48 percent ($\hat{d}$'=(a_2/a_1) comparing to the threshold level in the sample of non-advanced countries. This finding can be interpreted as evidence that advanced economies can sustain larger debt ratios than non-advanced economies.

The similar estimations for the effect of real long-term interest rates on fiscal policy are given by a_3 and a_4. The point estimate of a_3 is the impact of sovereign borrowing cost interacted with public debt, and it is statistically significant once the direct impact of interest rates and GDP-deflator inflation enter into the regression. Replicating our example using R7, the primary surplus would increase for countries with debt level is above a_4/a_3 (about 65 percent of GDP).

Table 4 and 5 show two substantial differences between the non-advanced and advanced economies based on the estimation results. First, the values of response coefficients in advanced countries are lower in magnitude than non-advanced countries. According to Mendoza and Ostry (2008), these results suggests that non-advanced economies have to converge to smaller mean debt-to-GDP ratios than advanced countries because the long-run expected value for debt-to-GDP ratio becomes lower in the

equilibrium when the size of response coefficient increases[12]. Therefore, the higher *p* in the sample of non-advanced countries does not indicate greater fiscal sustainability. It rather suggests that non-advanced countries require stronger fiscal policy responses of the primary surplus than advanced countries. Second, the coefficient values of output gap and expenditure gap in advanced economies are significantly larger than in non-advanced countries, which is in line with the findings of Mendoza and Ostry (2008). Hence, developing countries have less countercyclical fiscal policies, in turn; they have pro-cyclical primary balances. Mendoza and Ostry (2008) suggest that non-advanced economies should exploit debt markets more than they seem to do.

5.4. Time variation in Fiscal Response Function

In this section, we look at the time evolution of the "global" fiscal response function. The time variation in fiscal response function has also attracted a wide attention in academic literature. However, the interest was mainly restricted to single country analyses. This, in part, maybe due to obvious concerns about panel estimators. As in panels asymptotic distributions of estimators depend on relative convergence rates of N and T. Using panels; however, can give more valuable insights as global changes in solvency can potentially guide policy makers in individual countries to take adequate preventive measures. This is particularly relevant considering increasing globalization and complex interlinkages among world economies. For example, Russian's sovereign default in 1998 had severe consequences not only for countries that had tight trade or other (geographical, political, etc.) linkages with Russia, but also for countries like Brazil, Argentina and other emerging economies that are weakly related to it. Thus, it is important to understand how the global fiscal solvency changes overtime, specially, in presence of global shocks. We will see that the FE estimator completely wipes out effects of global shocks, while the CCEMG correctly accounts for global events and business cycles.

This section also guides us to understand "efficiencies" of FE and CCEMG estimators dealing with different types of biases when assumptions rendering them consistent and unbiased are violated. In fact, recursive and rolling estimations in panels are useful to better recognize the properties of different estimators. In this chapter, we further establish the importance of the CCEMG in macro-panel analyses, considering different types of possible biases in both estimators. Before presenting results and analyses, we briefly summarize them. Our concerns in the FE estimator are possible biases due to slope homogeneity assumption, and due to long time periods (or short N) in the data. The reason is that, FE is mostly micro-panel estimator which makes them consistent and unbiased when groups are homogenous in slopes, and N is larger than T. In CCEMG, the only concern is the possible bias because of short T. Our data consists of 61 consecutive years and 54 countries; however, unavailability of the

[12] See Bohn (2005) for a formal proof.

data for some groups in early periods drops the average T to 45. Thus, T is relatively long for the FE estimators, and may be short for the CCEMG. Thus, we try to find consequences of three main problems on the estimation FRFs: slope heterogeneity assumption in FE, small T in CCEMG, and finally large T in FE. To do that, we try to control for the number of periods (increasing T at each iteration), reducing the number of groups[13] (by dividing the sample into different sub-samples: advanced, non-advanced, mixed), by controlling for heterogeneity (this cannot be done completely in FE, but the heterogeneity can be reduced to some extent by considering sub-samples with similar characteristics, or/and by adding additional group specific covariates).

In FE we have two problems: short T (increasing T also, potentially, rises cross sectional dependence), and slope homogeneity assumption. Violations in both of them can cause biases that may affect the FRF not in the same direction. Our simulations give suggestive evidence that while slope heterogeneity in our sample causes FE to overestimate the FRFs; increasing time periods (increasing cross sectional dependence) causes underestimation. It is also likely that these two violations interact with each other in a more complex way that one needs to perform Monte Carlo simulations with the known data generating process.

In full sample, we find FRF's were higher in the FE estimator. When we reduce N, but also heterogeneity, at least hypothetically, (by dividing countries to advanced and developing countries), FE still overestimates the FRF for more than half of periods. However, when we reduce N, but still preserve heterogeneity (using mixed countries, inclusion is based on the availability of data), the overestimation is more exaggerated. In small, T there may appear underestimation with CCEMG, however we do not find strong evidence for that, or even if it exists, it disappears rapidly.

Moreover, in first three exercises there seems to be a convergence between two estimators. However, we conclude that this observation maybe misleading. We establish that, the convergence is likely due to two different biases that work in different directions. As a result, the evolution of FRF with FE estimator appears reverse "U" shaped. It is likely, that long T causes FE to underestimate the coefficient of lagged debt. Controlling further for heterogeneity to more extent (by adding more group-specific variables) and keeping T relatively long (T is set to 30 - to allow us to perform rolling estimations) shows substantial divergence in coefficients of the lagged debt with two estimators.

[13] Reducing N helps to avoid a bias caused by big N-T ratio (or small T bias) in the CCEMG; results still suggest higher FRF in FE.

5.4.1. Recursive Estimations

First, we recursively estimate simple FRF (namely, a model with lagged debt, output gap, expenditure gap) increasing the sample by one year at each step with starting window of 1950-1965.

Figure 4. FE and CCE estimators for lagged debt with 95% confidence intervals-Full sample

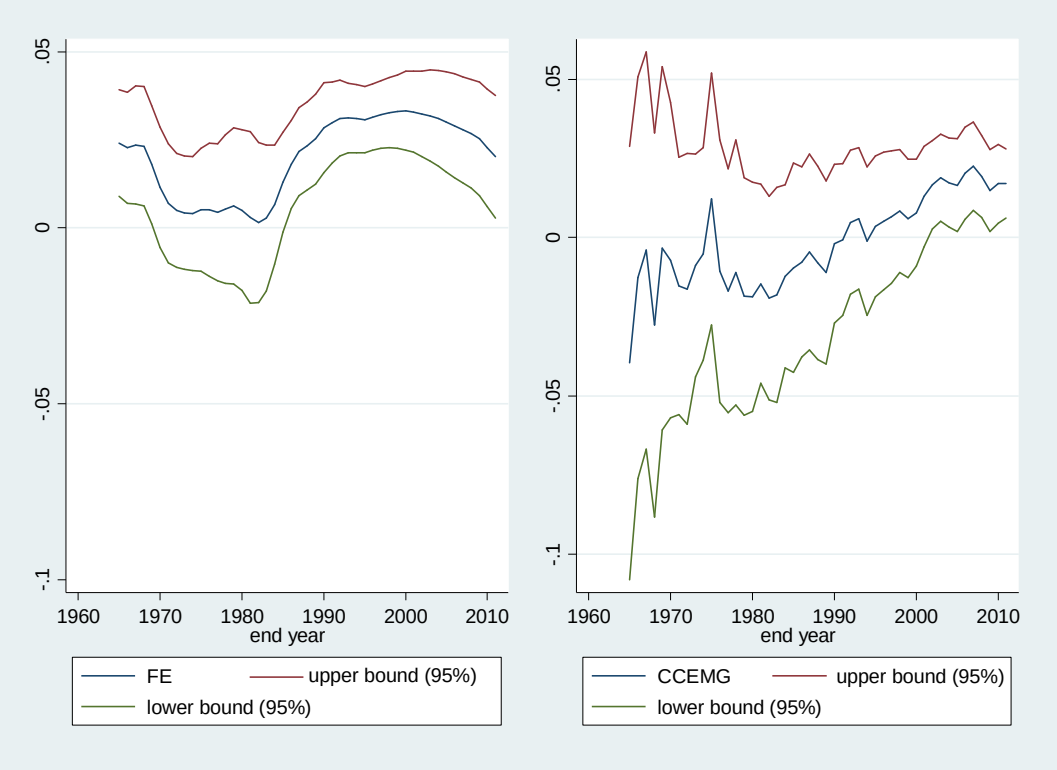


Figure 5. Fiscal Reaction functions

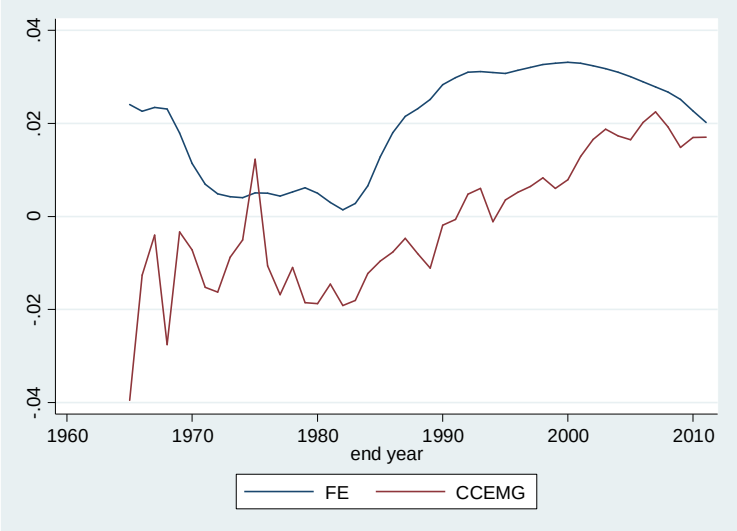


Figure 6. Kernel densities

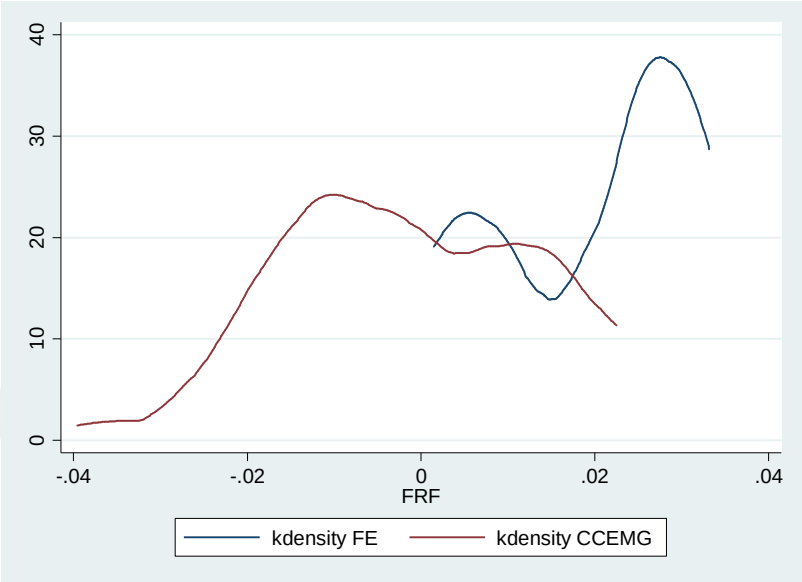


Figure 4 and 5 show that FE estimator finds substantial decrease in FRF in late 60's and early 70's, followed by almost a decade of low response to lagged debt. Then the response slowly increased, heading down in early 2000's – well before the global financial crisis. CCEMG estimator, however, starts with a negative response to the lagged debt where a large drop takes place in early and mid-70's, thus matching global energy crisis of 1973 timing wise[14]. The drop in FRF is relatively weak after the recent credit crunch, however, still pronounced.

Several apparent differences between two estimators are also worth some elaboration. First, standard deviations of CCEMG estimator decreases with T, while no obvious pattern is observed for standard deviations of FE estimator (Figure 4.). Second, FE (CCEMG) estimator "overestimates" ("underestimates") FRF for most of the time as shown in Figure 5.). As a result, the kernel density (Figure 6.) of the FE estimator is shifted to the right. Moreover, the time distribution of the FRF with FE is strictly positive and contracted, while for CCEMG it varies almost as twice as the one for the FE estimator (from -0.4 to 0.023), and is negative nearly 63% of the time (i.e. left skewed). However,

[14] However, we interpret large fluctuations at the beginning of simulations not entirely as a reaction to global shocks, but partly as a standard regression of the "initialization" bias and "small T" bias which is explained later

Figure 5. shows that the CCEMG estimates picks up positive values once observations starting from 1995 are included into the estimation.

First, results as expected considering limiting distribution properties of both estimates. While larger standard deviations with CCEMG estimator is mostly reflection of accounting for heterogeneous slope coefficients and is natural, differences in sizes of FRFs can be explained by examining asymptotic distribution of each estimator. For CCEMG, Chudik & Pesaran (2015) derive the asymptotic distribution of the estimate with lagged dependent variable and under weak exogeneity (including lagged values of weakly exogenous regressors). The original work by Pesaran (2006) abstracts from these assumptions, and for the derivation of asymptotic distribution of the estimator relative convergence rates of N and T are not restricted (as long as they jointly increase). However, in a dynamic panel model (such as the present which includes lagged debt), also possibly, with serially correlated factor loadings, the derivation of asymptotic distribution of CCEMG estimator requires $N/T \rightarrow x$ (where $0<x<\infty$)[15]. Therefore, when T is smaller than N the CCEMG estimator will be biased, and inconsistent. The paper is also accompanied by comprehensive use of Monte Carlo experiments where they analyse finite sample properties of CCEMG estimators and their bias-corrected counterparts in dynamic panels under numerous assumptions regarding the degree of cross-sectional dependence and parameter values. Their Monte Carlo experiments also extend to evaluate robustness of other estimates such as FE estimator, quasi-maximum likelihood estimator (QMLE) of Moon and Weidner (2015), and interactive effects estimator (IFE) of Bai (2009). Results of simulations suggest that in presence slope heterogeneity bias-uncorrected CCE estimators outperforms (in terms of the bias, power, and size) FE estimators, even when N and T is too small. FE estimator, however, could not only take into account the bias induced by slope heterogeneity, but also the time bias ($\mathbf{O}$ (T^{-1})) for T=40. As T increases the negative bias in autoregressive coefficient is dominated by the positive bias arising from slope heterogeneity resulting in over estimation by 0.07-0.10 (20-25% of the true coefficient). For the uncorrected version of CCEMG estimator the bias was around -0.12 when T = 40, but reduced dramatically to -0.02 for T=200.

Even though, Pesaran & Chudik (2015) is mostly concerned about biasness for the coefficient of lagged dependent variable, recursive estimation of the simple FRF suggests that the same concerns may equally apply for the coefficient of the lagged debt. In what follows we use terms "overestimation" and "underestimation" not in reference to the true unknown data generating process, but relative responses suggested by two estimators. We are also, in first place, are aware of the fact that underestimation with ($\mathbf{O}(T^{-1})$) bias in Pesaran & Chudik (2015) refers to the coefficient of the lagged dependent variable (known in literature as "Nickel" bias). Second, the theory does not give any guidance about the

[15] For the formal proof, see Chudik & Pesaran (2015)

direction of the bias. In fact, the sign of bias for the lagged debt can depend on the nature of heterogeneity (one being, e.g., its direction) or other underlying factors (such as assumptions regarding common factors; e.g. their numbers, heterogeneity in factor loadings etc.) which describe the true data generating process. Underestimation of the lagged independent variable with CCEMG may well be restricted only to our data set and cannot ubiquitously applied to other setups.

Moreover, keeping in mind results from Monte Carlo simulations of Chudik & Pesaran (2015), it is not completely clear weather FE estimator overestimates the coefficient of the lagged debt (because of the slope heterogeneity) or CCEMG underestimates it because of ($\mathbf{O}(T^{-1})$) bias. In fact, both may be the case. Therefore, before performing the same exercise for bias corrected version of CCEMG estimator, we would like not to rule out other possibilities. To be more precise, it is also possible that at the beginning of the sample countries were more heterogeneous in FRFs (suggesting larger standard deviations), and/or just the reaction was weaker (suggesting negative or smaller response coefficient).

To check for that, we split the data into two even subsamples across T dimension and estimate the FRF separately for each of them as shown in Table 6.

Table 6. Estimate of FRF for Two Even Subsamples

1950-1980		1981-2011	
FRF	SD	FRF	SD
-.019763	0186544	0280378 ***	0085856
CD		CD	
-2.19 [0.029]		1.12 [0.263]	

In fact, CCEMG estimator gives a negative and insignificant response coefficient with larger standard deviation for the first sub sample. Whereas, the coefficient is highly significant with smaller standard deviation for the second half of the sample. Cross-sectional independence cannot also be rejected for the post 1981 sample. These results suggest that while small sample bias can be present, there could well be other effects in action. Specifically, for the earlier sample, if we take into account the global energy shock, the Cold War, fierce rivalry between the capitalist world and several communist countries, many hot spots all over the world such as in Vietnam affected by the collision of interests of these opposing powers could well explain the weak response to the debt in presence of ever increasing need to profligate military and other expenditures.

Even though the above exercise may sound as a good explanation to divert attention from the “small T bias” of the CCEMG estimator, by no means, however, we find results in Table 6 sufficient on its own

to conclude that small T bias was not present. Examination of the estimation sample reveals that the average N/T ratio for the first sub-sample was 1.65, while it was 1.94 for the post 1981 data (the minimum N/T is higher in the first subsample, however)[16], indicating potentially strong evidence for ($\mathbf{O}(T^{-1})$) bias in both samples. Reporting minimum and average time periods in each estimation would yield relatively less information, as minimum T did not vary substantially, while average T naturally (by construction) increased as we did simulations increasing T at each step. Thus to seek for any further evidence for the relation between average N/T ratio and the reaction function, we apply the recursive estimation for different subsamples in N dimension, namely for advanced and developing countries in order to reduce N.

Below results (figure 7 and 8) are for advanced economics (N=24). For developed countries minimum observations for each group is 31, with 62 time periods. This should theoretically be sufficient to account for the ($\mathbf{O}$ (T^{-1})) bias in the full sample. The mean response function in CCEMG is positive for all recursions except for 1978, which is in contrast with our results for the full sample. At the start of the sample both estimators show huge drop in the FRF with substantially fluctuating CCEMG. This may be due to what we can call an "initialization bias" as when T is small adding new observations may change group averages rapidly, and because they are also included as regressors, these fluctuations are transmitted to the estimated FRF coefficients. Furthermore, FE starts to "overestimate" the response function from mid-80's. The average N/T ratio is smaller than unity consistently from the year 1975 and onwards, thus CCEMG should not have asymptotically "small T" bias, at least on theoretical grounds. Also, as T increases, FRF estimates change more smoothly. Differences between two estimators may also be due to the fact that when N is small, the asymptotic distribution of FE estimator is distorted as well, because it assumes and needs relatively large N to account for biases stemming from heterogeneous slope coefficients. Findings are in line with Monte Carlo experiments of Chudik and Pesaran (2015) (they find around 20-25% upward bias even N was relatively large - 200 groups) suggesting potential overestimation of the FRF by FE estimator.

Regarding the "economics" of time variation in FRF, CCEMG finds substantial differences over time how advanced economies reacted to the debt. The effect of Maastricht Treaty (1992) on fiscal responses for advanced economies sample (as the sample includes mainly European countries that were tightened together by the treaty) is strikingly clear. Whereas the response under FE started to increase almost a decade before the treaty was signed. We also observe an enormous decrease in solvency following the petrol crisis. For the FE estimator, however, effects of global shocks were largely smoothed out.

[16] Minimum N/T ratio is 4.77 and 4.60 for the first and the second subsamples, respectively.

Figure 7. FE and CCEMG estimators for FRF, with 95% confidence intervals – Advanced Economies

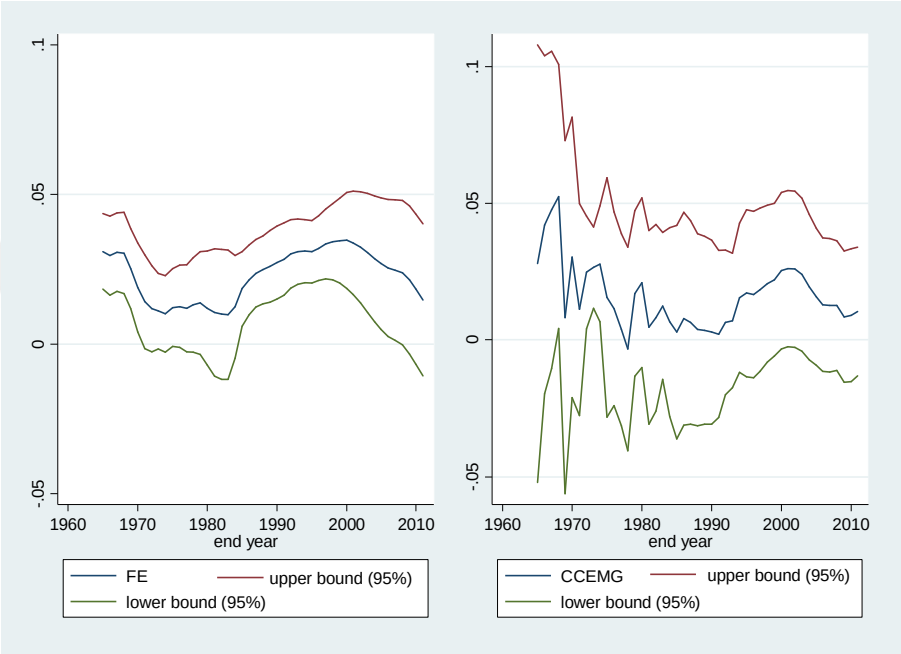


Figure 8. FE, CCEMG – Advances Economies

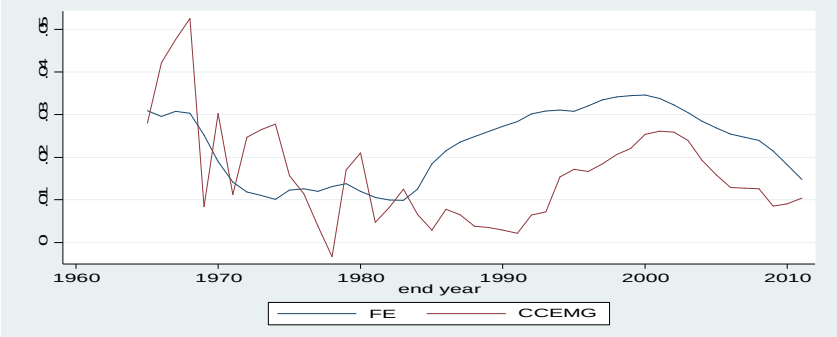
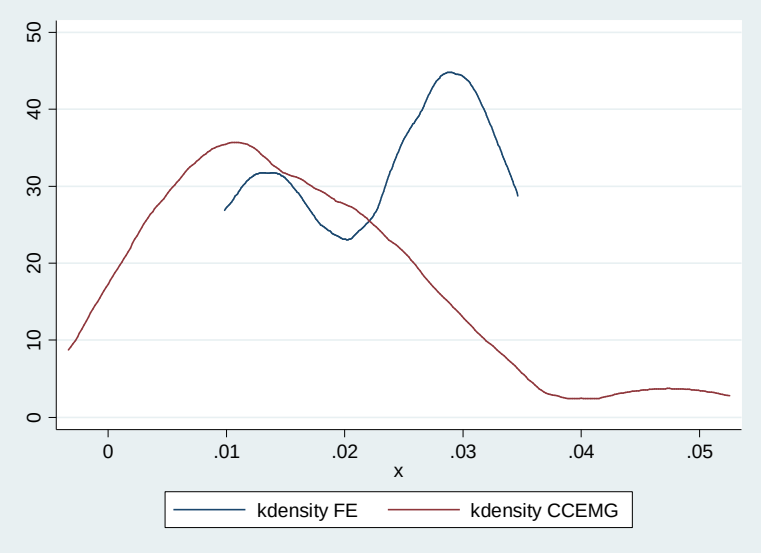


Figure 9. Kernel densities – Advanced Economies



We repeat the same exercise for developing countries. Due to limited data availability for emerging economies both the number of groups and the number of average time periods per a group are increasing as we roll over the sample, keeping N/T relatively constant (Figure 10) and below the unity for most of periods (40 out of 47). Thus, we can assume that (**O** (T^{-1})) bias is mostly accounted for in this sub-sample as well. Figure 11 shows that similar to advanced economies' sample both estimator give nearly similar results on average, while CCEMG estimator exhibits strong fluctuations at the beginning of sample again due to initialization bias[17]. Starting from 1980's, however these two estimates diverge until the end of 2000's. Second, both estimators find negative FRFs nearly entire first half of the sample. Most importantly, these results, give further insights for findings of the first exercise suggesting that negative FRFs of CCEMG estimator, till early 90's, for the full sample (advanced and non-advanced) simulations should not merely be attributed to (**O**(T^{-1})) bias, but largely are the result of negative coefficients implied by both estimates for developing countries. CCEMG estimator just picks up this heterogeneity between developed and emerging economies which is reflected in "full-N" simulations. Thus, instead the slope homogeneity assumption makes FEE to overestimate the FRFs.

[17] We also do not rule out that large variations in the beginning of sample could well be driven by true fluctuations in responses. In fact, bias corrected version of the CCEMG favours this impression.

We also find that the fiscal solvency of developing countries decreased largely after the recent financial crisis, while when we look at the results on advanced countries there is no substantial decrease in FRFs. For advanced economies, however, in the first half of 2000's, before the financial crisis, solvency decreased largely. This may be mere reflection of the "Great Moderation" that made advanced countries to be more optimistic and allowing them ignore developments in debt to GDP ratio. For developing countries, in contrast, the solvency consistently improved over time, if we ignore "global business cycles".

Figure 10. N/T ratio

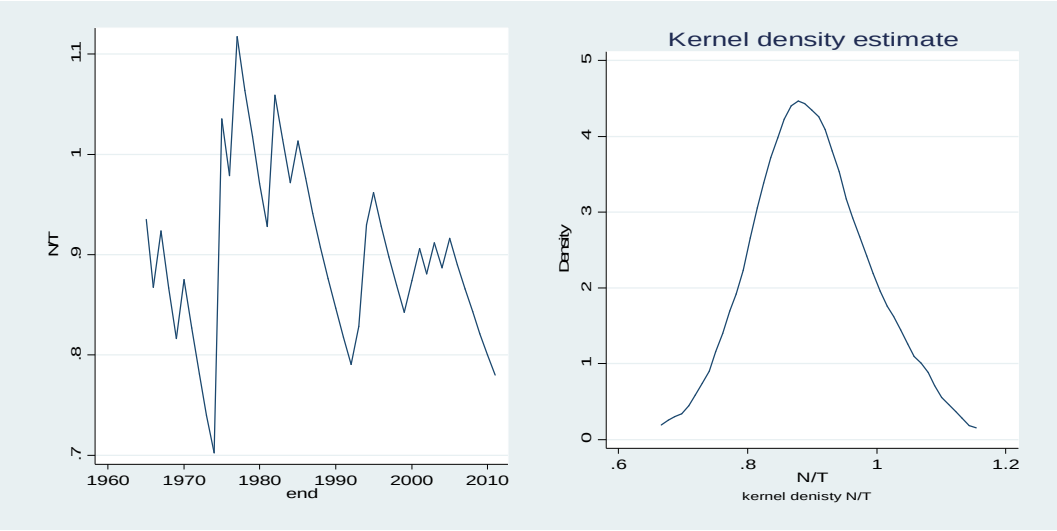


Figure 11. FE and CCEMG estimators for FRF, with 95% confidence intervals – Developing Economies

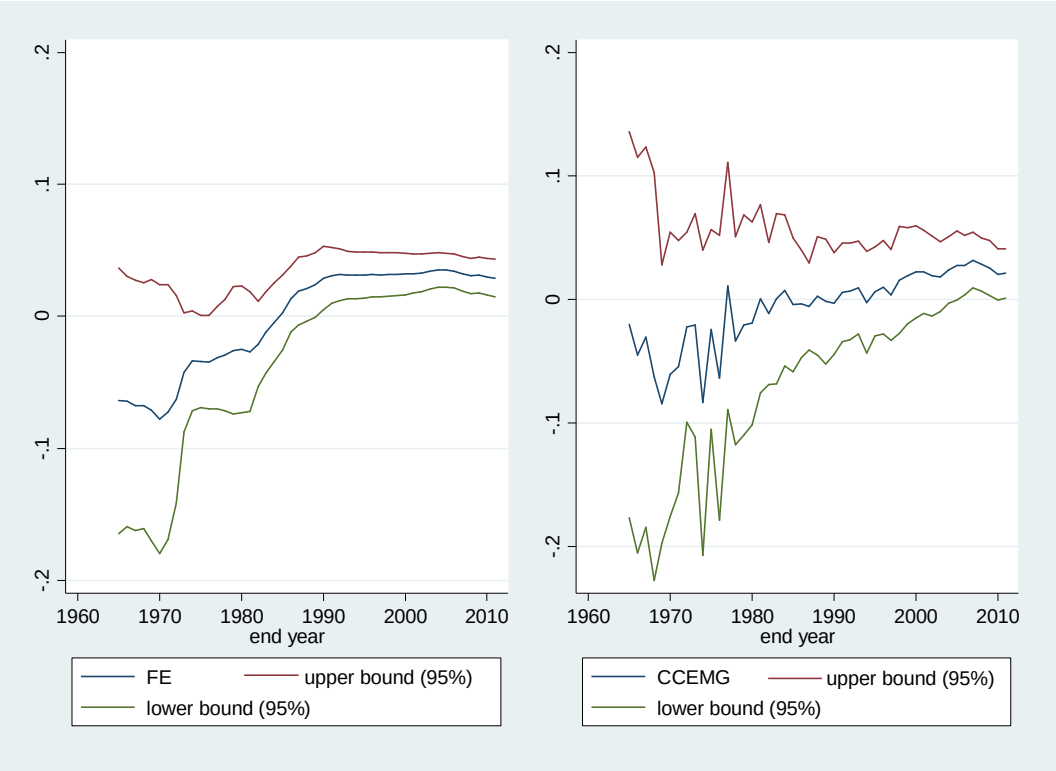
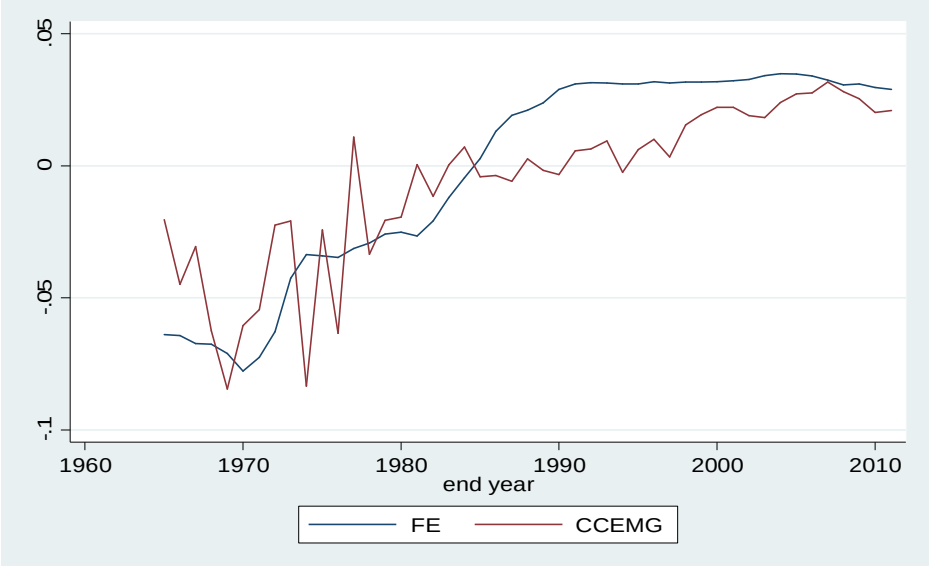


Figure 12. FE, CCEMG – Developing countries

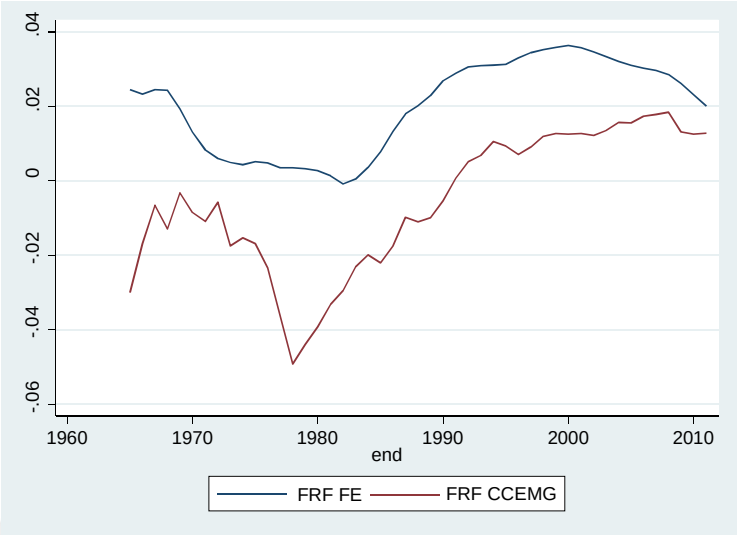


Although splitting the sample into advanced and non-advanced countries gives some useful insights, it still may work in "favour" of the FE estimator. The reason is that dividing the data by development potentially reduces the heterogeneity in sub-samples. To account for that, we reduce N, but using a sample that is mixed and includes both advanced and developing economies. Our selection criterion is purely based on the data availability in time dimension. We drop countries that have less than 50 observations on average for all four variables from developing countries, and less than 60 observations

for advanced economies[18]. This leaves us 35 groups (moderate N) where 13 of them belong to developing countries. As it shown in Figure 13, now the FE over estimates the fiscal response function consistently for all years.

[18] The threshold is bigger for the developed world because of the rich data availability in this group.

Figure 13. FE, CCEMG –Mixed Sample

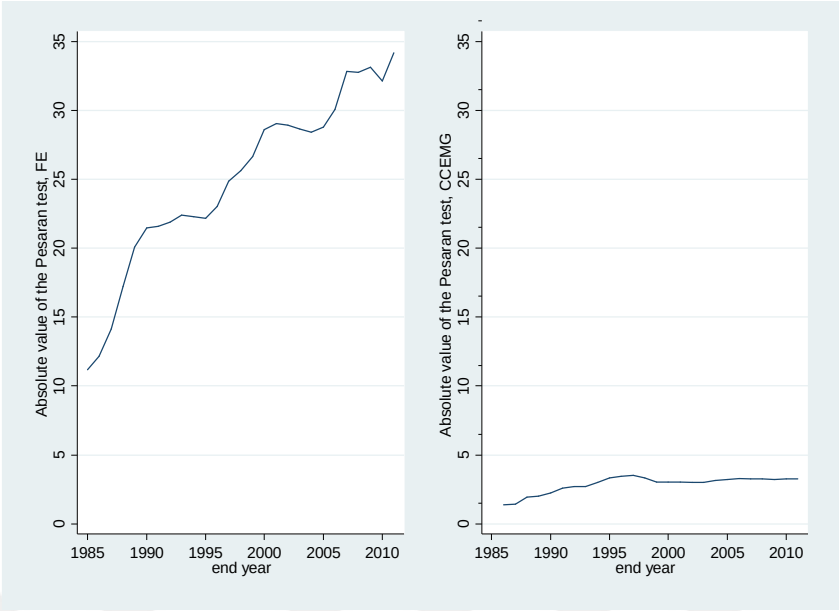


5.4.2 Rolling Estimations

We now draw the attention to another observation. All simulations above indicate nearly convergence in both estimators as we increase the number of years. It seems that for "full T" both estimators do not yield statistically different results. However, concluding the presence of "convergence" may be misleading. We did not mention yet assumptions regarding the asymptotic distribution of the FE estimator when T is not short anymore. Thus, we may consider another dimension of the potential bias by the FE estimators. As the asymptotic distribution of the FE estimator requires large N and short T, increasing number of year after some "threshold" can also induce a size bias that may or not work in the same direction with the bias introduced by slope heterogeneity. To understand the reason, we first establish the link between increasing T and cross-sectional dependence[19]. Figure 14 shows absolute values of Pesaran CD test when we increase T in our recursive estimations. For the FE estimator the cross-sectional dependence is high and consistently increases with the number of years. However, for the CCEMG estimator it is considerably low and relatively "stabilized" after 1995.

[19] Because the test cannot accommodate highly unbalanced data, we performed recursive estimations on advanced countries' data and with initial sample size of 1950-1985.

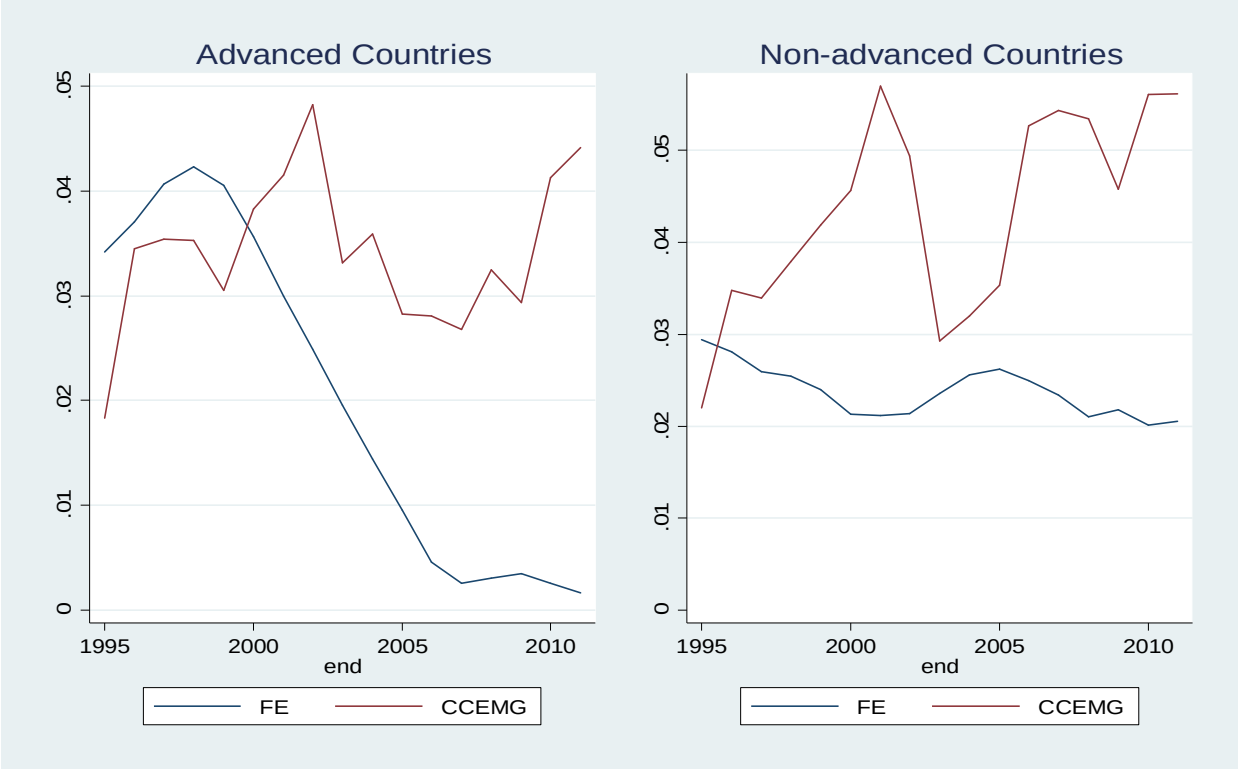
Figure 14 Absolute values of Pesaran cross-sectional dependence test



We can also check for the divergence of estimates, rolling the sample with constant T[20] in our simulations. To illustrate the point better we also attempt to control for the bias due to slope heterogeneity in the FE estimator to the extent we can. It allows us to detect better the marginal effect of "long T" while keeping heterogeneity bias relatively low. It is important as these two biases may work in opposite directions cancelling out each other. While it is impossible to fully account for heterogeneity with the FE estimator, we can reduce it to some extent by adding country specific variables, and country specific trends. Moreover, we apply simulations to advanced and non-advanced countries to further account for heterogeneity. The CCEMG estimator also includes the same set of covariates. Apart from country specific trends, additional country specific variables include growth-surprise and its' interaction with debt as they were the most available variables for almost all years and countries. Because we do not have enough observations on inflation and interest rate, they were excluded. Results of simulations suggest that, in fact, two estimators of the lagged debt diverged substantially, as suspected. The diversion is particularly strong for developed economies. The coefficient jumped from 0.002 to 0.044 for advanced economies, and from 0.020 to 0.056 for developing countries. These results suggest that the use of FE estimator with the inclusion of more country-specific variables underestimates the coefficient of lagged debt. In fact, this exercise shows "large T" bias in the FE estimator that needs to get more attention, at least as slope heterogeneity.

[20] We choose a moderate T, 30 years of window size. For the calculation of growth surprise, the window size is set at 15 to save more observations, but results are also robust to when it is set at 20.

Figure 15. Divergence of FE and CCEMG estimators.



To sum up, these exercises give further insights for potential problems in several dimensions (slope heterogeneity, cross sectional dependence) regarding the FE estimator that has been widely applied in previous works to analyse fiscal sustainability.

CONCLUSION

In the first part of the exercise, we search for the source of changes in fiscal solvency condition. Borrowing the dataset and model specifications from Mauro et al. (2015), we estimate panel regressions of primary balance responses to increases in government debt. To do this, we allow the government response to change depending on the deviations in the long-run growth rate and real long-term interest rate. We augmented their work by making use of a novel estimator (Common Correlated Effects Mean Group Estimator) which especially in our case addresses cross-sectional dependence and parameter heterogeneity.

The panel regressions are separated into three groups, namely, full sample, developing countries and developed countries. The results for the sample of all countries suggest that for countries with debt level excess of estimated threshold 27 percent of GDP (the majority of the countries), negative deviations from the long-run growth path have given rise to greater fiscal insolvency between the period 1950-2011. Mauro et al. (2015) find this threshold level as 9 percent of GDP in their corresponding regression for the same period. In comparison to our findings, their estimates suggest a much lower threshold debt level. The estimated debt threshold is higher (48 percent of GDP) for the advanced economies comparing to the level (23 percent of GDP) in the sample of non-advanced countries. This finding can be interpreted as evidence that advanced economies can sustain larger debt ratios than non-advanced economies. We also find that the interaction term of debt with long-term interest rate have a significant association with the size of the policy response to increasing debt level in advanced countries. In other words, increases in real long-term interest rate induce a stronger fiscal policy response to increasing debt-to-GDP ratio.

In the second part of the study, we take a closer look at the time evolution of the "global" fiscal reaction function. Even though, there can still be ($\mathbf{O}(T^{-1})$) with CCEMG, it seems to reduce rapidly with increased T. Furthermore, even if results from first three exercises (Figure 4, 5 and 6) seem to indicate a "convergence" between estimators, our last exercise (figure 14-15) suggests that it could be due to underestimation with FE in a large T data (due to increased cross sectional dependence). In fact, when we slightly control for heterogeneity and do rolling regressions these two estimators diverge considerably. Overall, simulations suggest that slope heterogeneity causes the FE to "overestimate" the FRF, but cross-sectional dependence (as it is increased with T) leads to a smaller response coefficient. Thus, the time evolution of the FRF is reverse "U" shaped with the FE.

We also want to stress that statements made in this section bear only positive character and only applies to the data set under study. To make normative and more reliable statements one needs to do these exercises with known data generating process, as direction of "biases" will depend on the nature and

the direction of heterogeneity, and many other underlying factors (e.g. assumption made on common factors and their evolution). However, at least, these exercises give useful insights how FE and CCEMG estimators of the FRF would differ depending on the time and country coverage of the data set.

APPENDIX

Table 1A 1. CD-statistics for the full sample using FE estimator

Primary Balance to GDP	R1	R2	R3	R4
1950-2011				
CD-Statistics	33.83	32.50	32.47	32.87
[p] value	[0.00]	[0.00]	[0.00]	[0.00]

Table 1A 2. CD-statistics for the sample of Non-Advanced Economies using FE estimator

Primary Balance to GDP	R1	R2	R3	R4
1950-2011				
CD-Statistics	11.66	10.88	10.92	10.80
[p] value	[0.00]	[0.00]	[0.00]	[0.00]

Table 1A 3. CD-statistics for the sample of Advanced Economies using FE estimator

pb	R1	R2	R3	R4	R5	R6	R7
1950-2011							
CD-Statistics	25.38	23.85	23.68	24.78	24.59	21.93	20.83
[p] value	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]	[0.00]

Stata Scripts

```
// Growth Surprise

//(save this as an ado file in the same folder)
program fcst_ar2, rclass
	syntax [if]

	qui tsset code year, yearly
			qui xtset code year
	local timevar `r(timevar)'
	local first `r(tmin)'

	regress l (0/2).rgc `if'
			local cf1= _b[_cons]
			local cf2= _b[l.rgc]
			local cf3= _b[l2.rgc]

			return scalar cf1=`cf1'
			return scalar cf2=`cf2'
			return scalar cf3=`cf3'

//(calculation of growth surprise variable)

cd z:\l14100\dosya\
use "son1.dta", clear
xtset code year
tsset code year
local winsize=20
local prjsize=5

rolling cf1=r(cf1) cf2=r(cf2) cf3=r(cf3), window(`winsize') recursive saving(ar2, replace): fcst_ar2

use "ar2.dta", clear
generate year=end+1
xtset code year
```

```
tsset code year

save "ar2.dta", replace

merge 1:1 code year using "son1.dta"
xtset code year
tsset code year

local winsize=20
local prjsize=5
local start_ =r(tmin)+`winsize'
local end_ =r(tmax)-`prjsize'-1



gen  growth_rate_f1=cf1 +cf2*L1.rgc+cf3*L2.rgc
gen  growth_rate_f2=cf1 +cf2*L1.growth_rate_f1+cf3*L1.rgc
gen  growth_rate_f3=cf1 +cf2*L1.growth_rate_f2+cf3*L2.growth_rate_f1
gen  growth_rate_f4=cf1 +cf2*L1.growth_rate_f3+cf3*L2.growth_rate_f2
gen  growth_rate_f5=cf1 +cf2*L1.growth_rate_f4+cf3*L2.growth_rate_f3



//gen lngdp=ln(realgdp)

//---> Static Projection of lngdp based on one step ahead growth rate forecast

 forvalues year = `start_'/`end_' {
 qui generate series`year' = .
 qui gen gr`year'=growth_rate_f1
 qui replace series`year'=lngdp if year<`year'
 qui replace series`year'=ln(1+gr`year'/100)+L.series`year' if year>=`year' & year<`year'+`prjsize'
 qui drop gr`year'

  }
 //---> Dynamic Projection of lngdp based on five step ahead growth rate forecast
 forvalues year = `start_'/`end_' {
 qui generate dseries`year' = .
```

```
  qui gen dgr1`year'=growth_rate_f1
  qui gen dgr2`year'=growth_rate_f2
  qui gen dgr3`year'=growth_rate_f3
  qui gen dgr4`year'=growth_rate_f4
  qui gen dgr5`year'=growth_rate_f5
  qui replace dseries`year'=lngdp if year<`year'
  qui replace dseries`year'=ln(1+dgr1`year'/100)+L.dseries`year' if year==`year'
  qui replace dseries`year'=ln(1+dgr2`year'/100)+L.dseries`year' if year==`year'+1
  qui replace dseries`year'=ln(1+dgr3`year'/100)+L.dseries`year' if year==`year'+2
  qui replace dseries`year'=ln(1+dgr4`year'/100)+L.dseries`year' if year==`year'+3
  qui replace dseries`year'=ln(1+dgr5`year'/100)+L.dseries`year' if year==`year'+4
  //remove temporary variables:
  qui drop dgr1`year'
  qui drop dgr2`year'
  qui drop dgr3`year'
  qui drop dgr4`year'
  qui drop dgr5`year'

  }

drop   growth_rate_f1
drop   growth_rate_f2
drop   growth_rate_f3
drop   growth_rate_f4
drop   growth_rate_f5

forvalues year = `start_'/`end_' {
   qui   tsfilter hp hpseries`year' = series`year', smooth(6.25) trend (trendhpseries`year')

}



forvalues year = `start_'/`end_' {
   qui   tsfilter hp dhpseries`year' = dseries`year', smooth(6.25) trend (dtrendhpseries`year')
```

```
}

tsfilter hp hplnrgdp = lngdp, smooth(6.25) trend(trendhplnrgdp)




generate grsp_static=.
generate grsp_dynamic=.

forvalues year = `start_'/`end_' {

     qui replace grsp_static=trendhplnrgdp-trendhpseries`year' if year==`year'
          qui replace grsp_dynamic=trendhplnrgdp-dtrendhpseries`year' if year==`year'
}

forvalues year = `start_'/`end_' {
          qui drop dhpseries`year'
          qui drop hpseries`year'
          qui drop dseries`year'
    qui drop trendhpseries`year'
          qui drop dtrendhpseries`year'
          qui drop series`year'

}
```

```
//Recursive Estimations

cd z:\l14100\dosya\
use " developing20to5.dta ", clear
xtset code year
tsset code year
generate lagd=l.d
```

```
tempname vector
      postfile `vector'  end                                    ///
                              beta_lagd_FE  ///
                              se_lagd_FE  ///
                              N_groups_FE  ///
                              using beta_FE_recursive, replace

set more off
drop if year<1950

local start_sample=1965

forv i = `start_sample'/2011 {
          display _n(16) in white _col(30) ///
          "Time period:"  `i'
   qui xtreg pb lagd  outputgap expgap    if year <=`i', fe robust
   post `vector' (`i') (_b[lagd])  (_se[lagd]) (e(N_g))
}
postclose `vector'
use beta_FE_recursive, replace



use " developing20to5.dta ", clear



xtset code year
tsset code year
generate lagd=l.d



tempname vector
      postfile `vector'  end                                    ///
                              beta_lagd_CCE          ///
                              se_lagd_CCE  /// standard devitation of beta
```

```
                        N_groups_CCE /// Number of groups for each year
                        using beta_CCE_recursive, replace


set more off
drop if year<1950

forv i = `start_sample'/2011 {
        display _n(16) in white _col(30) ///
        "Time period:"  `i'
  // qui xtdcce2 pb lagd outputgap expgap   if year <=`i', cr_lags(pb ) crosssectional(pb lagd
outputgap expgap )  jackknife
        qui xtmg pb lagd outputgap expgap  if year <=`i', cce robust
    post `vector' (`i') (_b[lagd]) (_se[lagd]) (e(N_g))
}
postclose `vector'
use beta_CCE_recursive, replace

        tsset end
        merge 1:1 _n using "beta_FE_recursive.dta"


twoway (kdensity beta_lagd_FE) (kdensity beta_lagd_CCE)
twoway (tsline beta_lagd_FE) (tsline beta_lagd_CCE)
gen down_FE=se_lagd_FE*(-1.96)+beta_lagd_FE
gen up_FE=se_lagd_FE*(1.96)+beta_lagd_FE
twoway (tsline beta_lagd_FE) (tsline up_FE) (tsline down_FE)
gen down_CCE=se_lagd_CCE*(-1.96)+beta_lagd_CCE
gen up_CCE=se_lagd_CCE*(1.96)+beta_lagd_CCE
twoway (tsline beta_lagd_CCE) (tsline up_CCE) (tsline down_CCE)
```

```
///Rolling Estimations


//---Divergence of Estimates---//
cd z:\l14100\dosya\
use " developing20to5.dta ", clear
```

```
xtset code year
tsset code year
generate lagd=l.d
tempname vector
	postfile `vector' start end                                    ///
                                beta_lagd_FE  ///
	    se_lagd_FE  ///
                                using beta_FE, replace

set more off
local window_size=30 //----set window size
local end_year_sample=2011
local start_year=1965 //---set start year: begining of the sample + number of years used to calculate growth surprise
local end_year=`end_year_sample'-`window_size'
	    forvalues i=`start_year'/`end_year' {
                local j=`i'+`window_size'
                display _n(16) in white _col(30) ///
                "Time period:" `i' "-" `j'
                 xtreg pb lagd outputgap expgap debtgrowth grsp_dynamic i.year  if year>=`i' & year<=`j' , fe robust
                  post `vector'                                        ///
                        (`i') (`j')                            ///
                        (_b[lagd])         ///
                        (_se[lagd])
                  local j=`j'+1

	}
	postclose `vector'

	set more on

	use beta_FE,clear
```

```
use " developing20to5.dta ", clear
xtset code year
tsset code year
generate lagd=l.d
tempname vector
      postfile `vector' start end                                   ///
                              beta_lagd_CCE          ///
          se_lagd_CCE   ///
                              using beta_CCE, replace

          set more off

          forvalues i=`start_year'/`end_year' {
                    local j=`i'+`window_size'
                    display _n(16) in white _col(30) ///
                    "Time period:" `i' "-" `j'
                     xtmg pb lagd outputgap expgap debtgrowth grsp_dynamic if year>=`i' &
year<=`j,', cce trend robust
                    post `vector'                                                ///
                              (`i') (`j')                              ///
                              (_b[lagd])        ///
                              (_se[lagd])
                     local j=`j'+1

      }
      postclose `vector'

      set more on

      use beta_CCE,clear

          tsset end
          merge 1:1 _n using "beta_FE.dta"
          twoway (kdensity beta_lagd_FE) (kdensity beta_lagd_CCE)
          twoway (tsline beta_lagd_FE) (tsline beta_lagd_CCE)
```

```
gen down_FE=se_lagd_FE*(-1.96)+beta_lagd_FE
gen up_FE=se_lagd_FE*(1.96)+beta_lagd_FE
twoway (tsline beta_lagd_FE) (tsline up_FE) (tsline down_FE)
gen down_CCE=se_lagd_CCE*(-1.96)+beta_lagd_CCE
gen up_CCE=se_lagd_CCE*(1.96)+beta_lagd_CCE
twoway (tsline beta_lagd_CCE) (tsline up_CCE) (tsline down_CCE)
```

REFERENCES

Abiad, A. & Ostry, J.D., 2005. Primary Surpluses and Sustainable Debt Levels in Emerging Market Countries. *IMF Policy Discussion Paper 05/6*, pp.1–19.

Afonso, A., 2005. Fiscal Sustainability: The Unpleasant European Case. *FinanzArchiv*, 61(1), pp.19–44.

Ahmed, S. & Rogers, J.H., 1995. Government budget deficits and trade deficits Are present value constraints satisfied in long-term data? *Journal of Monetary Economics*, 36(2), pp.351–374.

Baglioni, A. & Cherubini, U., 1993. Intertemporal budget constraint and public debt sustainability: the case of Italy. *Applied Economics*, 25(2), pp.275–283.

Berti, K. et al., 2016. Fiscal Reaction Functions for European Union Countries. *Discussion Papers, European Economy*, 28(April).

Blanchard, O. J. (1990). Suggestions for a New Set of Fiscal Indicators, OECD Economics Department Working Papers, No.79.

Bohn, H., 1995. The Sustainability of Budget Deficits in a Stochastic Economy. *Journal of Money, Credit and Banking*, 27(1), pp.257–71. Available at: http://ideas.repec.org/a/mcb/jmoncb/v27y1995i1p257-71.html.

Bohn, H., 1998. The Behavior of US Public Debt and Deficits. *Quarterly Journal of Economics*, 113(3), pp.949–963. Available at: http://www.mitpressjournals.org/doi/abs/10.1162/003355398555793.

Bohn, H., 2005. The sustainability of fiscal policy in the United States. *CESifo Working Paper*, (1446).

Bohn, H., 2007. Are stationarity and cointegration restrictions really necessary for the intertemporal budget constraint? *Journal of Monetary Economics*, 54(7), pp.1837–1847.

Bravo, A.B.S. & Silvestre, A.L., 2002. Intertemporal sustainability of fiscal policies: Some tests for European countries. *European Journal of Political Economy*, 18(3), pp.517–528.

Brock, W.A. & Durlauf, S.N., 2001. Discrete choice with social interactions. *Review of Economic Studies*, 68(2), pp.235–260.

Buiter, W.H. a., 1985. A Guide to Public Sector Debt and Deficits. *Economic Policy*, 1(1), p.113.

Callen, T. et al., 2003. Public Debt in Emerging Markets: Is it Too High? *World Economic Outlook: Public Debt in Emerging Markets*, pp.113–152.

Chudik, A. & Pesaran, M.H., 2015. Large Panel Data Models with Cross-Sectional Dependence: A Survey. *The Oxford Handbook Of Panel Data*, (May), pp.2–45.

Coakley, J., Fuertes, A.-M. & Smith, R., 2006. Unobserved heterogeneity in panel time series models. *Computational Statistics & Data Analysis*, 50(9), pp.2361–2380.

De Hoyos, R.E. & Sarafidis, V., 2006. Testing for cross-sectional dependence in panel-data models. *Stata Journal*, 6(4), pp.482–496.

Debrun, X. et al., 2008. Tied to the mast? National fiscal rules in the European Union. In *Economic Policy*. pp. 297–362.

Durlauf, S.N., Kourtellos, A. & Minkin, A., 2001. The local Solow growth model. *European Economic Review*, 45(4–6), pp.928–940.

Eberhardt, M. & Presbitero, A., 2013. This Time They Are Different: Heterogeneity and Nonlinearity in the Relationship Between Debt and Growth. *IMF Working Papers*, 13(248).

Eberhardt, M. & Teal, F., 2011. Econometrics for grumblers: A new look at the literature on cross-country growth empirics. *Journal of Economic Surveys*, 25(1), pp.109–155.

Eberhardt, M., 2012. Estimating panel time-series models with heterogeneous slopes. *Stata Journal*, 12(1), pp.61–71.

Enzo, C. & Ramon, V.H.J., 2003. Assesing Fiscal Sustainability: A Cross Country Comparison. *IMF Working Papers*, 145, pp.1–34.

Galí, J. et al., 2003. Fiscal Policy and Monetary Integration in Europe. *Economic Policy*, 18(37), pp.533–572. Available at: http://onlinelibrary.wiley.com.ezproxy.gsu.edu/doi/10.1111/1468-0327.00115_1/abstract\nhttp://onlinelibrary.wiley.com.ezproxy.gsu.edu/store/10.1111/1468-0327.00115_1/asset/1468-0327.00115_1.pdf?v=1&t=heic49ro&s=3e41abe54268680b8a51bb177e74b3d8d83f.

Ghosh, A.R. et al., 2013. Fiscal Fatigue, Fiscal Space and Debt Sustainability in Advanced Economies. *Economic Journal*, 123(566).

Ghosh, A.R. et al., 2013. Fiscal Fatigue, Fiscal Space and Debt Sustainability in Advanced Economies. *Economic Journal*, 123(566).

HAKKIO, C.S. & RUSH, M., 1991. IS THE BUDGET DEFICIT ???TOO LARGE???? *Economic Inquiry*, 29(3), pp.429–445.

Hamilton, J.D. & Flavin, M. a, 1986. On the Limitations of Government Borrowing : A Framework for Empirical Testing. *The American Economic Review*, 76(4), pp.808–819.

Juhl, T. & Lugovskyy, O., 2013. A test for slope heterogeneity in fixed effects models. *Econometric Reviews*, (785), pp.1–35. Available at: http://www.tandfonline.com/doi/abs/10.1080/07474938.2013.806708.

Krejdl, A., 2006. Fiscal Sustainability-Definition, Indicators and Assessment of Czech Public Finance Sustainability. *CNB Working Paper Series*, (3), pp.1–29.

Kremers, J.J.M., 1988. Long-run limits on the US Federal debt. *Economics Letters*, 28(3), pp.259–262.

Martin, G.M., 2000. US deficit sustainability: A new approach based on multiple endogenous breaks. *Journal of Applied Econometrics*, 15(1), pp.83–105.

Mauro, P. et al., 2015. A modern history of fiscal prudence and profligacy. *Journal of Monetary Economics*, 76, pp.55–70.

Mendoza, E.G. & Ostry, J.D., 2008. International evidence on fiscal solvency: Is fiscal policy "responsible"? *Journal of Monetary Economics*, 55(6), pp.1081–1093.

Payne, J.E., 1997. International evidence on the sustainability of budget deficits. *Applied Economics Letters*, 4(12), pp.775–779.

Pesaran, M.H. & Tosetti, E., 2011. Large panels with common factors and spatial correlation. *Journal of Econometrics*, 161(2), pp.182–202.

Pesaran, M.H., 2004. General Diagnostic Tests for Cross Section Dependence in Panels. *Iza*, (1240, (August)), pp.1–42. Available at: http://www.dspace.cam.ac.uk/handle/1810/446.

Pesaran, M.H., 2006. Estimation and inference in large heterogeneous panels with a multifactor error structure. *Econometrica*, 74(4), pp.967–1012.

Polito, V. & Wickens, M., 2005. Measuring Fiscal Sustainability. *Centre for Dynamic M Acroeconomic Analysis Conference P Apers*, 44(0), pp.1–56.

Quintos, C.E., 1995. Sustainability of the Deficit Process with Structural Shifts. *Journal of Business & Economic Statistics*, 13(4), pp.409–417. Available at: http://www.ncbi.nlm.nih.gov/pubmed/22115513.

Reinhart, C. & Rogoff, K., 2009. This Time It's Different: Eight Centuries of Financial Folly-Chapter1. *This Time It's Different: Eight Centuries of Financial Folly*, (17452), pp.2–20.

Reinhart, C. & Sbrancia, B., 2015. The Liquidation of Government Debt. *IMF Working Paper*, (WP/15/7), pp.1–47. Available at: http://www.imf.org/external/np/seminars/eng/2011/res2/pdf/crbs.pdf.

Smith, G.W. & Zin, S.E., 1991. Persistent deficits and the market value of government debt. *Journal of Applied Econometrics*, 6(1), pp.31–44. Available at: http://onlinelibrary.wiley.com/doi/10.1002/jae.3950060104/abstract.

Trehan, B. & Walsh, C.E., 1988. Common trends, the government's budget constraint, and revenue smoothing. *Journal of Economic Dynamics and Control*, 12(2–3), pp.425–444.

Trehan, B. & Walsh, C.E., 1991. Testing Intertemporal Budget Constraints: Theory and Applications to U.S. Federal Budget and Current Account Deficits. *Journal of Money, Credit, and Banking*,, 23(2), pp.206–223. Available at: http://www.jstor.org/stable/1992777.

Uctum, M. & Wickens, M., 2000. Debt and deficit ceilings, and sustainability of fiscal policies: An intertemporal analysis. *Oxford Bulletin of Economics & Statistics*, 2. Available at: http://eprints.whiterose.ac.uk/62714/.

Wilcox, D.W., 1989. The sustainability of government deficits: Implications of the present-value borrowing constraint. *Journal of Money, Credit and Banking*, 21(3), pp.291–306. Available at: http://www.jstor.org/stable/1992415.

World Economic Outlook database (http://www.imf.org/external/pubs/ft/weo/2014/02/weodata/index.aspx)