UNIVERSITY OF TAMPERE

UNIVERSITÉ DE FRIBOURG
UNIVERSITÄT FREIBURG

Master’s Degree Programme in Public Economics and Public Finance (MGE)

Master’s Thesis

The effects of fiscal policy on long-term interest rates Evidence from 29 OECD countries

Under the supervision of:

Prof. Hannu Laurila
School of Management, University of Tampere

Prof. Dr. Thierry Madiès
Department of Political Economy, University of Fribourg

Kristian Tötterman

Tampere, October 2017

Kalevantie 4
33100 Tampere

Totterman.Kristian.K@student.uta.fi
+358 50 3575163

Abstract

University of Tampere and University of Fribourg

Master's Degree Programme in Public Economic and Public Finance (MGE)

TÖTTERMAN, KRISTIAN: The effects of fiscal policy on long-term interest rates; Evidence from 29 OECD countries

Master's Thesis: 62 pages, 3 appendix pages

October 2017

Keywords: fiscal policy, budget deficit, public debt, long-term interest rates, OECD countries, panel data, fixed effects estimation, omitted factors, financial crisis of 2007–08

How does fiscal policy affect long-term interest rates? Despite the broad literature in this field of research, the results on both the magnitude of the effect and under which conditions it prevails still remain vague today. This thesis seeks to clarify these issues. It begins by discussing the link between fiscal policy and interest rates from a theoretical point of view and by reviewing the empirical literature on the topic. The empirical part of the thesis then applies fixed effects estimation on a panel of 29 OECD economies over the last three decades. A so-called baseline regression with common specifications is estimated first, as it provides the base case for our empirical analysis. The model is then expanded in several ways to tackle factors that are omitted in the baseline regression but have strong theoretical reasoning to be included in the model.

We find the baseline regression to describe poorly the variation of long-term interest rates with the post-financial crisis data, but to provide plausible results when it is estimated solely for the pre-crisis data. Our results imply that this is due to certain complexities caused by the crisis that cannot be tackled with the baseline regression alone. By extending the model, the presence of unconventional monetary policy tools, credit ratings and capital account openness are all found to have an important effect on the relationship between fiscal policy and long-term interest rates. Furthermore, our results imply that the addition of the prior two measures successfully tackles most of the complexities caused by the crisis. Hence, we propose them to be an important part of a robust setup for studying the interest rate effect of fiscal policy with post-financial crisis data.

As has been typically found in the prior literature, our results show that fiscal policy comes over to long-term interest rates mainly through the flow variable. While a one percentage point deterioration in the primary balance-to-GDP ratio raises long-term interest rates by 13 basis points, an equivalent change in the public debt-to-GDP ratio has an interest rate effect of only one basis point at the most, which is not statistically significant. Additionally, we find the weak interest rate effect of the public debt-to-GDP ratio to decline even further as we add credit ratings to the model. This implies that bulk of the effect of the stock variable on long-term interest rates comes via the default premium, and once sovereign creditworthiness is controlled for with another measure, the public debt-to-GDP ratio becomes irrelevant for long-term interest rates.

Tiivistelmä

Tampereen yliopisto ja Fribourgin yliopisto

Master’s Degree Programme in Public Economics and Public Finance (MGE)

TÖTTERMAN, KRISTIAN: The effects of fiscal policy on long-term interest rates; Evidence from 29 OECD countries

Pro Gradu -tutkielma: 62 sivua, 3 liitesivua

Lokakuu 2017

Avainsanat: finanssipolitiikka, budjettialijäämä, julkinen velka, pitkät korot, OECD-maat, paneeliaineisto, kiinteiden vaikutusten malli, huomioimattomat tekijät, finanssikriisi 2007–08

Kuinka finanssipolitiikka vaikuttaa talouden korkotasoon? Huolimatta siitä, että kysymystä on tutkittu laajalti, ei finanssipolitiikan korkovaikutuksesta tai sen suuruudesta ole muodostunut yleistä konventiota ekonomistien keskuudessa. Tämä tutkielma pyrkii valaisemaan finanssipolitiikan korkovaikutusta. Työn alussa käydään läpi finanssipolitiikan vaikutuskanavat valtionlainojen pitkiin korkoihin sekä tehdään lyhyt katselmus aiheen empiiriseen kirjallisuuteen. Tutkielman empiirisessä osiossa sovelletaan kiinteiden vaikutusten mallia 29 OECD-maan paneeli-aineistoon vuosilta 1989–2016. Ensin estimoidaan nk. yleinen regressio, joka noudattaa kirjallisuudelle tyypillistä spesifikaatiota ja tarjoaa siten hyvän vertailukohdan kirjallisuuden muihin tutkimuksiin. Tämän jälkeen tarkastelua laajennetaan niin, että yleiseen regressioon lisätään useita teorian valossa tärkeitä muuttujia, jotka eivät tule alkuperäisessä mallissa huomioiduksi.

Tulosten perusteella yleisen regression voidaan todeta kuvaavan heikosti valtionlainojen pitkien korkojen vaihtelua finanssikriisin jälkeisellä aineistolla mutta tuottavan uskottavia tuloksia kriisiä edeltävällä aineistolla. Tulokset vihjaavatkin siten, että yleinen regressio on riittämätön finanssipolitiikan korkovaikutuksen tutkimiseen kriisin jälkeisellä aineistolla. Laajentamalla mallia löydetään epätavanomaisen rahapolitiikan, luottoluokitusten sekä pääomataseen avoimuuden vaikuttavan tilastollisesti merkitsevästi finanssipolitiikan ja pitkien korkojen väliseen yhteyteen. Kahden aiemman muuttujan sisällyttäminen malliin näyttää lisäksi korjaavan suurimman osan kriisin tuomista estimointiongelmista. Tältä pohjalta ehdotankin, että niiden huomioiminen on tärkeää kestävien tulosten saavuttamiseksi, kun finanssipolitiikan korkovaikutusta tutkitaan finanssikriisin jälkeisellä aineistolla.

Finanssipolitiikka näyttää välittyvän pitkiin korkoihin pääasiassa budjettialijäämien kautta. Yhden prosentin suuruinen budjettialijäämä nostaa pitkiä korkoja 13 korkopisteellä, kun taas vastaavan suuruinen muutos maan julkisen velan BKT-osuudessa muuttaa korkotasoa enintään yhdellä korkopisteellä, eikä vaikutus ole tilastollisesti merkitsevä. Velkatasomuuttujan heikko korkovaikutus katoaa lisäksi lopullisesti, kun valtioiden luottoluokitukset lisätään malliin. Tämän perusteella näyttää siltä, että julkisen velan BKT-osuuden heikko korkovaikutus syntyy pääasiassa luottoriskikanavan kautta, ja kun valtion luottokelpoisuus otetaan huomioon suoremmin kuin maan julkisen velan BKT-osuuden avulla, tulee velkatasosta valtionlainojen pitkien korkojen kannalta merkityksetön.

Contents

List of figures

List of tables

List of appendices

List of abbreviations

ADF	Augmented Dickey-Fuller test
ADV	Advanced countries
AIC	Akaike information criterion
CDS	Credit default swap
DGLS	Dynamic generalized least squares
GIPSI	Greece, Ireland, Portugal, Spain and Italy (also known as GIIPS or PIIGS)
GMM	Generalized method of moments
G10	Group of Ten
EC	European Commission
EME	Emerging market economies
EMU	Economic and Monetary Union
EO	Economic Outlook
EU	European Union
FAP	Factor augmented panel
FE	Fixed effects
IFS	International Financial Statistics
IMF	International Monetary Fund
IPS	Im-Pesaran-Shin test
ISO	International Organization for Standardization
LM	Lagrange Multiplier test
OECD	The Organisation for Economic Co-operation and Development
OLG	Overlapping generations model
OLS	Ordinary least squares
POLS	Pooled ordinary least squares
RE	Random effects
SCP	Stability and Convergence Programme
S&P	Standard & Poor’s
UMP	Unconventional monetary policy tool
VAR	Vector autoregression
2SLS	Two-stage least squares
3SLS	Three-stage least squares

1 Introduction

Governments that have access to international capital market can effectively finance additional investment or expenditure by debt issuance. While this brings flexibility to budget planning, it also adds up to governments' interest payments and so reduces the funds disposable for all future spending. The total economic cost of fiscal expansion does not equal the mere nominal interest payments though, but instead depends on various different implications that public debt has on the economy. For the sake of the sustainability of public finances, the interest rate effect of fiscal policy has a particular role. If increased fiscal laxity also increases interest rates of the economy, then the government's costs increase, not only by the interest payments on the additional stock of debt, but also by the higher refinancing costs on the entire stock of debt. The interest rate effect of fiscal policy may also have other, more indirect, adverse implications on the economy: higher interest rates can, for example: reduce investment, restrain spending on interest-sensitive durable consumption and cut down consumption via a negative wealth effect[1]. While it is difficult to measure the total economic cost of the above effects, it is easy to see that it depends directly on the degree to which fiscal expansion actually raises interest rates. (Engen & Hubbard, 83, 2004)

From 1990 to 2017 the average public debt-to-GDP ratio among OECD economies effectively doubled. The majority of this development took place during two periods of time: the first half of the nineties, and the Great Recession triggered by the financial crisis of 2007–08. While the deterioration of fiscal balances during the first half of nineties occurred in only roughly ten countries, the financial crisis shook public-sector finances in almost all the OECD members. (OECD, 2016) The overall growth of public debt is shocking by itself, yet even more worrying are its adverse implications on the sustainability of public finances in the developed countries. By pointing out the harsh economic development of the GIPSI countries (Greece, Italy, Portugal, Spain and Ireland) since the wake of the crisis, one could easily claim the injuriousness of large-scale fiscal deterioration and how the bulk of its effect is due to higher interest rates that threaten economies' fiscal sustainability. Although the case of the GIPSI countries makes a fine example of pick-ups in interest rates due to fiscal deterioration, it largely lacks generalization.

[1] Rising of interest rates decreases the real value of assets held by households and firms, hence it discourages consumption.

This thesis seeks for a more general answer. It studies the interest rate effect of fiscal policy, assesses its magnitude and the conditions under which it prevails, as the evidence on all of these issues still remains vague today despite the broad literature in this field of research (Aisen & Hauner, 2013, 2501). In order to answer these questions, we identify the different channels through which fiscal policy can be expected to influence interest rates and briefly review the empirical literature, aiming at building up the necessary knowledge for carrying out the empirical assessment of the interest rate effect, which is the main objective of the thesis. In the empirical part of the thesis we apply fixed effects estimation on a real-time panel of 29 OECD economies over the last three decades. We first estimate a so-called baseline regression that is most suitable for cross-study comparison due to common specifications as to provide the base case for our analysis. We then extend the model in several ways as to tackle omitted factors that have strong justification to be included in the model from theoretical point of view.

We contribute to the literature by applying the baseline regression on a dataset that reaches the post-financial crisis era, which is not covered in the prior literature apart from the paper by Dell'Erba and Sola (2016). We find that the results of the baseline regression are problematic with our full sample and contradict with many of the results found in the prior literature. Our evidence suggests that this is due to certain complexities caused by the financial crisis of 2007–08 that cannot be dealt with by the baseline regression alone. By extending the model, we find the presence of unconventional monetary policy tools, credit ratings and capital account openness to have an important effect on the relationship between fiscal policy and long-term interest rates. Furthermore, we find that majority of the issues due to the crisis are successfully tackled with the addition of the prior two measures, which we propose to be an important part of a robust setup for studying the interest rate effect with post-financial crisis data. Under this specification, fiscal policy comes over to long-term interest rates mainly through the flow variable, i.e. the primary balance-to-GDP ratio, and the effect through the stock variable, i.e. the public debt-to-GDP ratio, is much weaker, just as found in the prior literature.

The remainder of the thesis is structured as follows. Section 2 discusses the link between fiscal policy and interest rates from a theoretical point of view while section 3 briefly reviews the recent empirical literature on the topic. Section 4 presents the research data, carries out the model selection as well as studies the time-series properties of the data. Section 5 goes through the empirical results, discusses their relevance and compares them to the findings of the other studies in this field of research. Section 6 concludes.

2 The link between fiscal policy and interest rates

2.1 Main channel of impact

From among the different channels through which fiscal policy could affect interest rates the one occurring via national savings can be considered the main one (Baldacci & Kumar, 2010, 3). The economic theory on this matter, however, proves inconclusive because contradictory views exist on this channel. Particularly, Ricardian equivalence proposition suggests highly different results for debt-financing government spending than the conventional view of debt[2]. We briefly cover both views in turn.

In short, the conventional analysis suggests that budget deficits crowd out national savings, which in turn hinder capital formation of an economy. In the long run, this leads to a smaller stock of capital yielding a higher marginal product of capital, or interest rate. The simplest framework that generates this outcome is the classical one with Solow's (1956) growth model. Consider the following change: a government decides to finance some of its spending by issuing bonds instead of levying taxes and leaves the amount of spending untouched. In other words, it shifts from a balanced budget to a budget deficit, and so develops a positive stock of debt. As the government budget balance also denotes public savings, such a shift represents a fall in public savings for the full amount of the deficit. The deficit also increases consumers' current disposable income by an equal amount, as the new means of financing government spending imply less (current[3]) taxes to consumers. This additional income is then devoted to consumption and private savings based on the marginal propensity of each action (which is assumed to be strictly positive but smaller than one for both actions). Therefore, as a result of the budget deficit, private savings increase less than one-to-one with disposable income, implying an overall fall in national savings. This is the idea of budget deficits crowding out national savings in all simplicity, and once it is brought into the Solow's growth model it results

[2] By conventional view of debt, we refer to the one suggested by Elmendorf and Mankiw (1999). The implications of government debt according to this view are based on the assumption that the economy is Keynesian in the short run and classical in the long run. The name stems from the belief by the authors that this is in fact the predominant view. (Elmendorf and Mankiw, 1999, 1627–1630)

[3] Within this framework, consumers are not forward-looking and so the consumption decisions depend solely on their current disposable income.

in a balanced equilibrium growth path with a lower capital stock yielding a higher interest rate. (For the full process in detail, see e.g. Mankiw, 2012, Ch. 3 and 9)

This is the usual starting point whenever this channel of impact is examined. It leaves many unanswered questions though, such as: Would the interest rate response to a budget deficit be any different if consumers were forward-looking, saving and consumption decisions in each period were determined as a solution to an intertemporal optimization problem, or external debt was used as a source of funding instead of internal debt. Therefore, taking a quick look at a more formal examination that tackles these issues is worthwhile. Despite being more than five decades old, Diamond's (1965) seminal paper provides a good example of such an examination. Diamond (1965) is fully credited for the following analysis and so will not be further referenced.

A neoclassical growth model with overlapping generations (OLG) consists of utility maximizing individuals who live for two periods, working in the first while being retired in the second. Individuals receive earnings only in the period they work, which they allocate to consumption over both periods of their lives. To accomplish this, they channel part of their earnings as one-period loans to entrepreneurs who seek to employ capital for production. The working individuals make up the supply side of the capital market, leaving the demand side to entrepreneurs. Thus, the number of savings in each period is determined by a competitive market process, where individuals ultimately set the amount of savings by solving their intertemporal consumer problem. As the market is competitive, the entrepreneurs employ any amount of capital the individuals decide to save, paying them its marginal product, or interest rate. In the long run, the model converges into a steady-state equilibrium, where all key variables, such as savings, capital-labour ratio and interest rate, remain constant over time, as is typical for a growth model.

To find out the economic implications of fiscal policy, government is then added to the original model. In this framework, the government only has one-off spending in the initial period, which it finances fully by issuing one-period bonds. Instead of repaying the debt fully when it comes due, the government rolls it over to the next period by issuing new debt so that it keeps the debt-labour ratio constant over time. In other words, it issues new debt so that the growth rate of nominal government debt equals that of labour. The long-run equilibrium interest rate arising when permanent and stable debt-labour ratio exists is then compared to that of the original model.

The long-run equilibrium interest rate raises when public debt exists because of the taxes needed to cover part of the interest cost of government debt. As just described, the government is automatically able to finance the interest cost up to the amount of the growth rate of labour simply by issuing new debt. The remainder of the cost, which is the difference between the interest rate and the growth rate, must be dealt with by collecting taxes from individuals[4]. These taxes affect the capital supply, as they imply less income to individuals in all but the initial period (tax collection begins only in the second period). This can be seen as a negative wealth shock to all future generations, and as consumption in both periods of individuals' lives is a normal good, the taxes reduce their savings. In the long run, decreased savings again lead to a smaller stock of capital yielding higher interest rate, just like in the standard approach. This effect takes place regardless of the source of funding, and is in fact the only effect to occur if government borrowing is externally funded. However, if internal funding is used instead, a second effect can be expected to take place. This is because internal debt also alters the demand side of the capital market as some of the domestic savings are channelled to non-productive government bonds instead of productive physical capital. This further reduces the capital stock of the economy and increases its marginal productivity, or interest rate.

Despite the small differences in the pass-through mechanism from the budget deficit to the equilibrium interest rate, both approaches arrive at the same conclusion that fiscal deterioration ultimately causes the equilibrium interest rate to rise as less savings are channelled to productive capital investments (assuming the case of dynamically efficient equilibrium in Diamond's (1965) model). One important thing to note is that both approaches treat budget deficits as permanent and so indirectly imply that it is in fact the cumulative amount of budget deficits, i.e. the stock of public debt, which affects the level of equilibrium interest rate. There are also studies, like the book by Auerbach and Kotlikoff (1987), which consider the interest rate effects of temporary changes in fiscal balances. These will, however, not be further discussed in the thesis for two reasons. First, the short-run effects of budget deficits on interest

[4] It has to be pointed out that in an economy with an equilibrium interest rate smaller than the growth rate of labour, this implies negative taxes to individuals. In this case, the interest rate effect of public debt is negative. This is a dynamically inefficient equilibrium though, since consumption and so utility in each period could be increased simply by lending out less to entrepreneurs. (Diamond, 1965, 1137)

rates, which are the only ones that can be expected to take place due to temporary budget deficits, are likely to be trivial compared to the long-run effects (Auerbach & Kotlikoff, 1987, 89). Second, the actual development of fiscal balances in OECD economies during the last three decades matches better to the assumption that the budget deficits are mostly permanent[5].

Ricardian equivalence proposition provides an alternative viewpoint to the economic implications, or rather, dis-implications, of public debt. In short, it states that incurring public debt has no effect on national savings, interest rates nor on any other macroeconomic variable. This is because if a government is unable to run a Ponzi game, i.e. the present value of its debt cannot be positive in the limit (Romer, 2012, 586–590), then a tax reduction brought about by a budget deficit must be joined by an equal-sized increase in taxation in the future. In other words, this means that the budget deficit does not reduce the tax burden on individuals but merely postpones it. Rational and forward-looking individuals anticipate this and alter their current savings in order to meet the future tax liability. As a result, private savings increase just by one-to-one with the fall in public savings (the budget deficit), leaving national savings, along with capital stock and interest rate, unchanged.[6] (Elmendorf & Mankiw, 1999, 1641) In fact, the only thing to change under the Ricardian view is the decomposition of the national savings.

Robert Barro's 1974 paper revived the academic debate on the relevance of the Ricardian proposition. He studied the economic implications of public debt in a similar OLG framework that was developed by Diamond (1965) with one central difference: in his model, the utility of individuals depends not only on their own consumption but also on the utility of their descendants. This intergenerational altruism ends up connecting all the different generations into a single network and thus makes the consumer problem very much like that of infinitely living households (Bernheim, 1989, 63). Needless to say, Barro's simple modification to the model results it yielding no real effects from budget deficits, hence supporting the Ricardian view.

[5] Out of the 29 OECD economies observed, only Belgium, Denmark, Netherlands and New Zealand managed to decrease the debt-to-GDP ratio from 1990 to 2017 (OECD, 2016).
[6] This analysis is conditional to the assumption that the budget deficit does not entail any changes in government spending, and if this assumption is relaxed, the a priori conclusion can no longer be guaranteed (Elmendorf and Mankiw, 1999, 1641).

Generally speaking, one could say that the Ricardian view is heavily reliant on two key assumptions: the government budget constraint (the so-called No-Ponzi-Game condition) and the permanent income hypothesis[7] (Elmendorf & Mankiw, 1999, 1641), and that the models which comply with these assumptions yield a supporting result for the theorem. The dependence on the strong assumptions is, however, the very same reason why the Ricardian view is often dismissed among the economic practitioners. In addition to the two assumptions above, Bernheim (1989, 63) lists seven more that it takes for the Ricardian result to hold. These are:

> *1) successive generations are linked by altruistically motivated transfers;*
> *2) capital markets are either perfect, or fail in specific ways;*
> *3) consumers are rational and farsighted;*
> *4) the postponement of taxes does not redistribute resources across families with systematically different marginal propensities to consume;*
> *5) taxes are non-distortionary;*
> *6) the use of deficits cannot create value (not even through bubbles); and*
> *7) the availability of deficit financing as a fiscal instrument does not alter the political process.*

As one can undermine the Ricardian result by relaxing any of these assumptions (or either of the two mentioned before), it is not a surprise that the view is dismissed by the majority of economists.

Figure 1 provides an illustration of how the conventional view of debt and the Ricardian equivalence proposition differ in their views on how fiscal deterioration affects interest rates (and other key macroeconomic variables, such as national savings, output and exchange rate). It posits the economic implications of public debt from the point of view of United States as the domestic economy. Under the conventional view, the implications are further split into two alternatives based on whether the source of funding for the deficit is solely external or at least partly domestic.

[7] The permanent income hypothesis states that individuals' consumption decisions are based on their lifetime income rather than their current income (see Friedman (1957)).

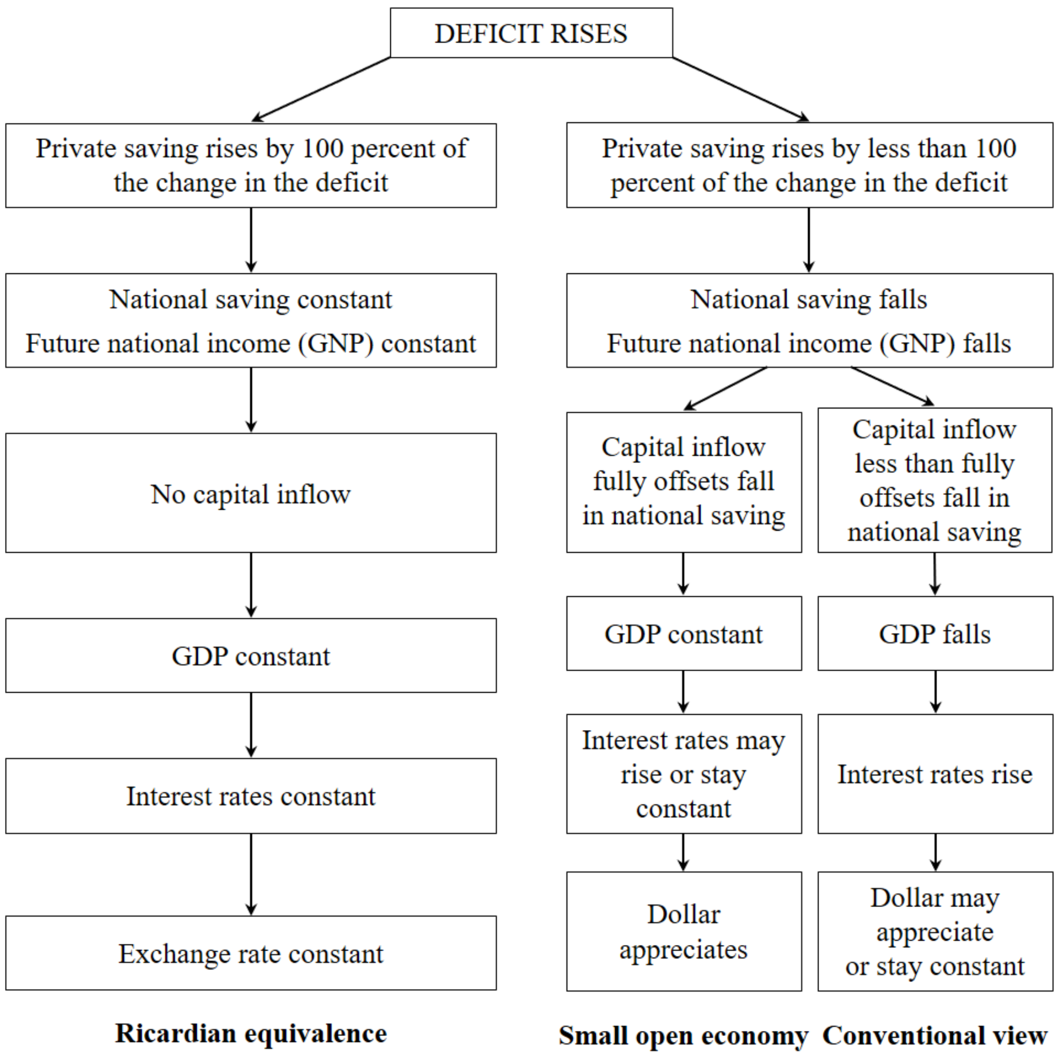


Figure 1. Theoretical responses to a change in the budget deficit (Gale and Orszag, 2004, 113).

The left branch of figure 1 shows how under the Ricardian view the decreased public savings brought about by the budget deficit are fully offset by increased private savings and as a result, all the other variables remain constant. Under the conventional view, which is shown on the right, the response in private savings is smaller and hence fiscal deterioration has an effect on national savings, interest rates and other macroeconomic variables. The magnitude of the effect depends on, among other things, the source of funding for the deficit so that the smaller the externally funded share of the deficit, the larger its expected impact. For a small open economy in a world with perfect capital mobility, this could mean that all other macroeconomic variables remained constant, except for the national savings and the exchange rate (middle branch of figure 1). This has barely been covered here, but will be looked more closely in subsection 2.3.

Despite we expect to find that a positive impact from fiscal deterioration on interest rates, and in fact focus more on how to best formulate this relationship, it is fair to point out that there exists not only an alternative but also an opposing view on the interest rate effect of fiscal policy[8]. In addition, the bequest motive embodied in the modern Ricardian view cannot be ignored to influence the saving behaviour of individuals, although it seems implausible that it would alone make the revenue side of fiscal policy totally irrelevant. Nonetheless, it could very well be that the bequest motive dampens the crowding out effect of fiscal policy and so makes the interest rate effect of budget deficits smaller than is expected by the conventional view.

2.2 Effect through risk premiums

As interest rates are determined in capital market where suppliers of capital meet up with demanders of capital, taking note of the factors that affect investor behaviour is crucial. Capital asset pricing model states that the expected return on any asset must equal the risk-free return in the capital market plus the risk adjustment. Fiscal policy could hence affect interest rates too by increasing a few risk premiums embodied in the nominal yields of government bonds. We take note of three possible risk premiums that could be affected by domestic fiscal measures: default, inflation and liquidity premium.

If adverse developments in fiscal balances imply to investors that the default risk of the government has increased, they will demand more yield for holding the government's debt securities in their portfolios. It remains unclear, however, to what extent do declines in fiscal measures actually increase the default risk of governments. Japan, for instance, has the highest burden of public debt (ca 234 % of GDP) and has had the largest government primary deficits (average primary deficit of 4.4 %) among all the OECD economies over the last 28 years, and yet it has a credit rating of A+ from S&P. As a comparison, Poland, whose two fiscal policy measures are way better than Japan's (ca 68 % and 1.6 %, respectively), possesses a credit rating two notches worse from S&P than Japan does. (OECD, 2016 and S&P Global Ratings, 2017) Needless to say, there are many other important factors to influence the default risk of governments, and so by looking at merely fiscal policy measures one cannot jump to any conclusions. Standard & Poor's, for example, performs institutional, economic, external, fiscal

[8] In fact, there are many empirical studies that have examined the interest rate effect of fiscal policy and found support for the Ricardian equivalence (see e.g. Plosser (1982, 1987), Evans (1987a, 1987b) and Boothe and Reid (1989)). This further implies that the Ricardian view should not be totally disdained.

and monetary assessment of the economy when determining its sovereign's creditworthiness (S&P Global Ratings, 2014).

Credit ratings are only one way to measure sovereign risk. Other, more market-oriented ways to measure sovereign risk are e.g. the long-term government bond spreads and the credit default swap (CDS) spreads (Beirne & Fratzscher, 2013, 61). While the former denotes the gap between yields of two sovereign debt securities with similar tenors, the latter is the difference of market values of two insurance contracts against the default of debtor sovereigns. All of these three measures have their own disadvantages. Credit ratings, for example, are discrete in nature and so are much less sensitive to any new information than the two other measures. On the other hand, the value of the two spread measures may also be affected by various other risk premiums which makes their interpretation harder. Despite the differences, the three measures are in fact highly correlated. (Beirne & Fratzscher, 2013, 66–72) Figure 2 plots the average prices of 5-year CDS contracts from the last quarter of 2016 against the respective credit ratings to illustrate the relationship between these two variables for 23 OECD economies which for the CDS prices are available on Bloomberg. ISO 3166-1 Alpha-2 codes denote the country names.

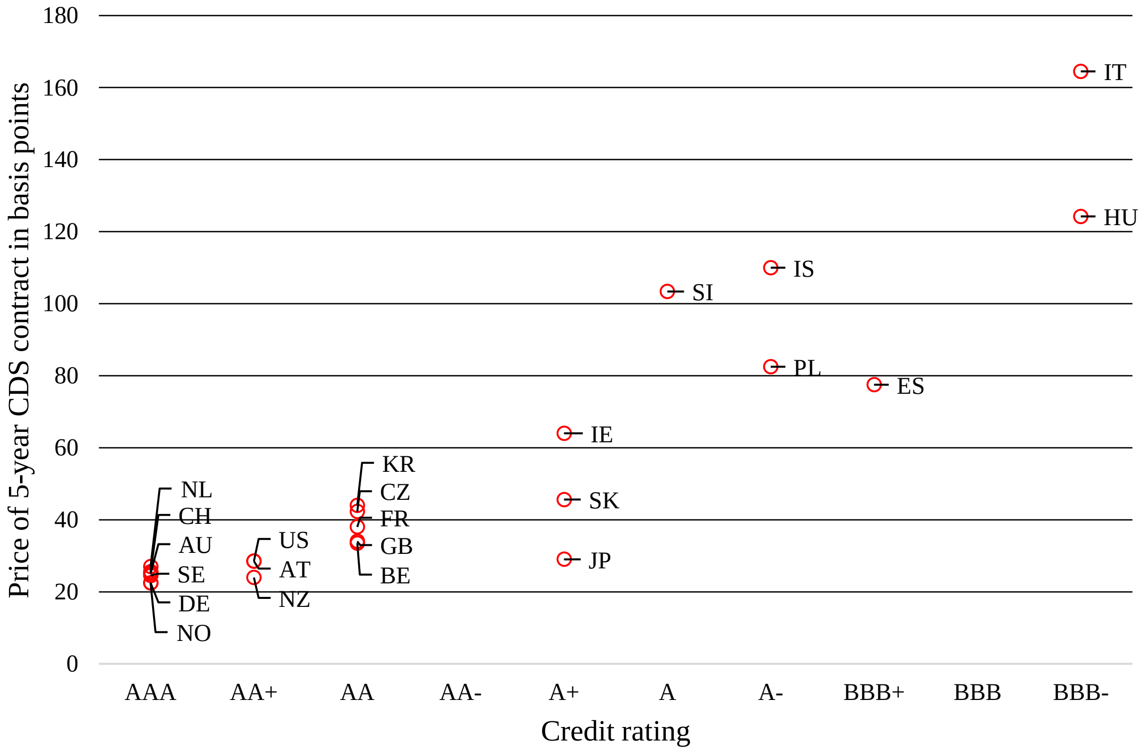


Figure 2. CDS contract prices and credit ratings (Bloomberg and S&P Global Ratings, 2017).

Although the degree to which declines in fiscal balances raise the default risk of governments is unclear, there is empirical evidence that the measures, particularly the market-oriented ones, are responsive to changes in domestic fiscal conditions. For example, Beirne and Fratzscher (2013) found that both government bond yield and CDS spreads depend statistically significantly on public debt-to-GDP and fiscal balance-to-GDP ratios. Aizenman et al. (2013) found similar evidence by employing only CDS spreads, showing that they depend on public debt-to-tax base ratio and fiscal balance-to-tax base ratio. The sensitivity of these measures to domestic fiscal conditions seems to have increased after the financial crisis, especially among the GIPSI countries, indicating that severe mispricing of sovereign risk took place in financial markets during the pre-crisis period 2000-07 (Beirne & Fratzscher, 2013, 60 & 71). Given the rating downgrades since the wake of the crisis, one can reach a similar conclusion regarding the credit ratings of sovereigns (Arezki et al., 2011, 7). Although the different measures of default risk seem not to have been entirely reliable during all times, investors and credit rating agencies certainly take domestic fiscal conditions into account when pricing the default risk of sovereigns.

Additionally, domestic fiscal conditions are identified to be important for predicting and classifying sovereign debt crises. In fact, Manasse and Roubini (2009) rank public external debt-to-fiscal revenue ratio to be the second most important determinant of sovereign default, second only to total external debt-to-GDP ratio. In addition, the type of sovereign debt crisis associated to unsustainable levels of public debt is the one where governments face insolvency, which further underlines the importance of this link (governments can also face e.g. illiquidity or various other macroeconomic risks in a sovereign debt crisis). (Manasse & Roubini, 2009, 201)

Just as demanding higher compensation for lending to governments more likely to default, rational investors also look for higher yields to offset the inflation erosion in their nominal returns. Hence, if declines in fiscal balances increase inflationary pressure in an economy, domestic interest rates are likely to increase along with the rate of expected inflation. Fiscal deficits could be expected to cause inflationary pressure in an economy for at least two reasons: for producing or enlarging positive output gaps and for raising concerns about monetization of debt (Baldacci & Kumar, 2010, 4).

In a simple IS-MP-IA[9] framework, the immediate effects of (permanent) expansionary fiscal policy are a higher level of output and interest rate at any given level of inflation[10]. As inflation responses to developments in real economy only gradually over time, it stays on its initial level at first and starts to pick up slowly only if the expansionary fiscal policy brought the level of output above its natural rate. If it did, however, inflation targeting central bank would increase the real interest rate in order to prevent the inflation from accelerating to excess. Gradually rising inflation and real interest rate would then cause the economy to adjust towards its new long-run equilibrium, where both real interest rate and inflation remain steady at a higher level and output again equals its natural rate. (Romer, 2000) The important point for our story is that the IS-MP-IA model suggests that both the real interest rate and inflation will pick up in the long run due to permanent budget deficits.

On the other hand, if large and persistent budget deficits indicated the unsustainability of fiscal policy, investors could begin to distrust the ability of government to pay off its debt without inflating it away, and arguably this mere concern could raise current inflation. In principle, a government has two ways to finance its deficits: it can either print high-powered money or issue debt in capital market. If the government runs continuous deficits and chooses the latter option to finance them all, its stock of public debt grows continuously as well[11]. It is likely, however, that there exists an upper limit in the demand for government bonds in capital market, and hence the public debt cannot grow infinitely. This means that a government running this kind of unsustainable fiscal policy scheme eventually must monetize its debt, which will lead to an increase in the price level in a monetarist economy. If investors anticipated this sequence of events by looking at current and expected future deficits, inflation would pick up immediately instead of rising at the time of expansion of monetary base. (Sargent & Wallace, 1981) This is one of the reasons why the independence of central banks is so extensively

[9] IS-MP-IA is used as illustrative framework as it captures the relationship between demand shocks and inflation more intuitively as does the traditional IS-LM-AS framework (Romer, 2000, 157).

[10] The immediate response of interest rate actually depends on whether the MP curve is modelled to respond solely to inflation or to both inflation and output, and is either zero or positive, respectively. As the model arrives to the same long-run conclusion with either assumption, the choice between the two is irrelevant. (Romer, 2000)

[11] It is crucial to assume that the demand for government bonds is such that the bonds pay higher interest rate than is the growth rate of economy. Otherwise the debt-to-GDP ratio will decline over time and hence it is sufficient for the government to finance its deficits fully by debt issuance in capital market. (Sargent & Wallace, 1981)

demanded, as political pressures would most likely cause an inflationary bias to monetary policy decisions (Mishkin, 2004, 352).

Despite both business cycle and monetarist theories suggesting that persistent fiscal deficits are inflationary in the long run, the data has rarely been found to support this argument very strongly (Catão & Terrones, 2005, 529). Some recent evidence exists though. For instance, Catão and Terrones (2005) find a strong and robust positive relationship between fiscal deficits and inflation rates, which seems to prevail not only during high and exceptionally high rates of inflation but also during moderate rates of inflation. Lin and Chu (2013) arrive to a very similar conclusion, further stressing out the importance of nonlinearity in this relationship; persistent fiscal deficits seem to be more inflationary in high- and middle-inflation economies than in low-inflation economies.

Liquidity of an asset is another factor which affects the yield that investors demand for holding it in their portfolios. Simply put, it describes the ease of converting the asset into cash at need. Liquidity is a desired attribute, especially against market stress events, where the need to liquidate even large sums of assets may arise in no time. In this type of situation assets that trade often in the market and offer a good price certainty, i.e. have better liquidity (Mishkin, 2004, 125), outperform assets without these types of qualities. Investors, being risk-averse on average, are willing to hold illiquid assets only for the compensation, or liquidity premium, that they offer over liquid assets.

Typically, liquidity premium is used to describe why long-term bonds offer higher yield over short-term bonds from the same issuer, i.e. why the yield curve is upward-sloping. Both liquidity premium and preferred habitat theories explain this phenomenon (although using somewhat different approaches) by stating that due to the higher interest rate risk in long-term bonds the only reason investors are willing to hold them over short-term bonds is in exchange for higher return. This results to an upward-sloping yield curve even in the absence of expected interest rate hikes. (Mishkin, 2004, 133-134)

Liquidity premium can also be used to explain why two bonds with similar qualities from different issuers can have different expected returns: Bonds from one issuer, which are more illiquid, can be expected to offer higher return over more liquid bonds from the other issuer. This is where fiscal positions come into the picture. The absolute amount of public debt determines the size of the sovereign bond market and so can be expected to affect the liquidity of the government bonds. Market size can be considered as a measure of market depth[12]: Larger markets may indicate lower information costs since securities are likely to trade more frequently and a larger investor base analyses their features (Gómez-Puig, 2006). This makes finding a buyer on the fly an easier task than in an illiquid market.

There is some evidence to support the idea of absolute amount of public debt affecting market liquidity and so liquidity premium, although not very recent results exist to the knowledge of the author. Inoue (1999) notes that the stock of public debt has a positive effect on the market liquidity in G10 countries while McCauley and Remolona (2000) state that although the relationship between size and liquidity is complicated by several factors, size matters for the liquidity in a case where important fixed information costs about the future path of interest rates exist. According to their estimation there might be a threshold of around $100-200 billions, which it takes to sustain a liquid government market. In addition, Gómez-Puig (2006) finds that relative market size levels could have positively impacted the widening of adjusted spreads within 9 EMU countries from 1996 to 2001. He notes this effect to be strongest among small economies of the sample and having accentuated with EMU, indicating further that the relationship between outstanding stock of public debt and liquidity premium is nonlinear.

2.3 Other affecting factors

We note two other factors that may play an important role when the link between budget deficits and interest rates is examined. These are openness of the economy and the stance of monetary policy that economy chooses to employ. We briefly cover both in turn.

As already noted in subsection 2.1, the source of funding can be crucial in determining how strongly persistent fiscal deficits affect domestic interest rates. This is because external debt does not have the second effect of crowding out productive private capital that internal debt has, and because of this it can be expected to have a lesser impact on interest rates (Diamond, 1965). Hence, one could conclude that at given deficit, the larger is the externally funded share of it, the smaller is its expected interest rate impact. At extreme, this could mean that if external

[12] Market depth is one of the three traditional measurements of market liquidity. It describes the volume of transactions that can be executed without impacting the price of the asset. (Bervas, 2006)

public debt was allowed to grow unlimitedly and used as the only source to fund fiscal deficits, the productive capital stock (interest rate) of the economy would not have to decrease (increase) at all due to decreased national savings. In fact, this is the result that the conventional view suggests for a small open economy, in a world with perfect capital mobility (See e.g. Mankiw, 2012, Ch. 6). It is worth to note that even if this was a realistic outcome for a small open economy, it would likely not come without a cost, as building up a negative net foreign investment position would also mean an equal-sized cumulative trade deficit for the economy, as well as appreciation of its currency. Furthermore, it is unclear to what extent it would be sustainable to cover continuous fiscal deficits by acquiring larger and larger negative net foreign investment position. Over time, this could very well still raise domestic interest rates, yet through a different channel. (Baldacci & Kumar, 2010, 4-5)

Baldacci and Kumar (2010) find indirectly supporting evidence that the degree of openness of the economy affects the interest rate impact of fiscal deficits by showing that in economies where deficits are mostly domestically financed, larger effects on interest rates can be expected to take place. Aisen and Hauner (2008) investigate this more directly and find that both the amount of domestic financing and the degree of capital account openness of the economy largely affect how fiscal deficits come over to domestic interest rates: they estimate the coefficients for high domestic financing and low capital account openness to be 92 and 67 basis points, respectively (Aisen & Hauner, 2013, 2507). Both findings by Baldacci and Kumar (2010) and Aisen and Hauner (2013) support the result achieved by Dellas et al. (2005), who find that large trade and capital mobility (high degree of openness and capital mobility, respectively) empower fiscal policy, even though they find neither of these effects to be statistically significant (Dellas et al., 2005 ,15).

The stance of monetary policy is an obvious factor that needs to be addressed whenever the interest rate effect of fiscal policy is being studied. Central bank, being the monopoly supplier of high-powered money, can significantly affect the prevailing interest rates in an economy by setting the price at which it is willing to deal with the money markets[13]. The quantitative effect of a change in this price, or rate, on other interest rates, depends, among other things, on the term of the interest rate in question: the chosen policy rate is most closely connected to short-

[13] Despite central banks differ in their monetary policy implementation methods, this is the general idea of how they control interest rates.

term money-market rates but the link weakens as the term of the interest rate lengthens. (Bank of England, 1999, 4) Expectations theory explains this phenomenon by stating that long-term rates are mere averages of current and expected future short-term rates, and thus changes in current short-term rates play a smaller role at the long end of the yield curve (as expected future short-term rates make up a larger share of the average) (Mishkin, 2004, 129). Expected future short-term rates are again closely related to expected future policy rates, which further depend on current expectations about future levels of real economy factors in the interest of the central bank, such as inflation, employment and output. Hence, at the long end of the yield curve, more room is left for expectations on real economy factors, whereas at the short end, the interest rates are firmly anchored by the chosen policy rate of the central bank. Therefore, we can also expect other factors, like fiscal policy, to have a larger impact on long rather than short-term interest rates of the economy.

Although long-term interest rates are expected to be less affected by the current stance of monetary policy, short-term interest rates, which, again, are closely related to the chosen policy rate of the central bank, should not be ignored when the interest rate effect of fiscal policy is being studied. This is due to two reasons. First, as long-term interest rates equal the geometric mean of current and expected short-term interest rates plus the term premium, the prevailing short-term interest rates surely affect their long-term counterparts via the term structure of interest rates: The higher the short-term interest rates, the higher their long-term counterparts, other things equal. The second factor is more indirect and in fact very closely related to the impact of expected inflation on interest rates. In addition to causing an upward pressure on the overall yield curve, higher inflation expectations can be expected to raise expected short-term interest rates if investors expect the central bank to tackle the inflation upsurge by raising future policy rates. A set of expectations like this is likely to steepen the yield curve, raising the spread between long-term and short-term interest rates. This is the reason why e.g. Canzoneri et al. (2002) and Faini (2006) rationalise that the interest rate effect of fiscal policy should mainly affect the spread between short and long-term interest rates rather than interest rates of any particular tenor. For the reasons stated above, many researchers, like Ardagna et al. (2007), Baldacci and Kumar (2010), Clayes et al. (2012), Dell'Erba and Sola (2016), Faini (2006) and Gruber and Kamin (2012), take the short-term interest rates into account in addition to inflation expectations when studying the interest rate effect of fiscal policy in order to control the cyclical conditions, stance of monetary policy and the prevailing term structure of interest rates.

Besides setting the policy rate, central banks can influence interest rates also by other methods such as forward guidance, quantitative easing, credit easing and liquidity injections. These types of methods have become common tools in the arsenal of all leading central banks during the last decade when they were faced with increasing challenges of tackling threat of deflation and market dislocations. (Pill & Reichlin, 2017, 1) Due to having being unusual before the post-financial crisis period, these methods are still commonly referred to as unconventional monetary policy tools (UMPs).

The details of UMPs, both with respect to their implementation and possible impacts, vary largely across the different central banks along with their exact monetary policy objectives (Fawley & Neely, 2013, 52), which makes their careful assessment a burdensome task. Luckily, we suffice with the information that one of the main reasons to employ UMPs is the need from central banks to conduct expansionary monetary policy when short-term interest rates are at or close to their effective lower bound and hence cannot be further lowered by adjusting the policy rate (Joyce et al., 2012, F272). Utilizing short-term interest rates alone to address the stance of monetary policy when the effective lower bound is binding could cause misspecification of the stance of monetary policy, as the short-term interest rates would likely to fail to fully take into account the employed stance. This is the reason why we consider the impact of the presence of UMPs on long-term interest rates, although a more careful assessment of their effects is left for the monetary policy literature.

2.4 A brief conclusion of affecting factors

This subsection aims at providing a summary of the theoretical framework discussed so far. We denote the yield on 10-year government bond by R_t, the yield on 1-year government bill by r_t, the expectation formed at time t by E_t and the risk premium by θ_t[14]:

$$R_t = \frac{1}{1+N}\left(r_t + \sum_{i=1}^{N} E_t r_{t+i}\right) + \theta_t$$

[14] We follow the notation by Mankiw and Miron (1986).

In effect, the above equation divides the yield on 10-year government bond into two sub factors, the average of the current and expected future short-term interest rates $\frac{1}{1+N}(r_t + \sum_{i=1}^{N} E_t r_{t+i})$ and the risk premium θ_t, just as is posited by the expectations theory. With this type of lay out, we are able to study how deterioration in fiscal balances can be expected to affect the two sub factors of long-term interest rates through the different channels of impact. Figure 3 provides an illustration of this framework.

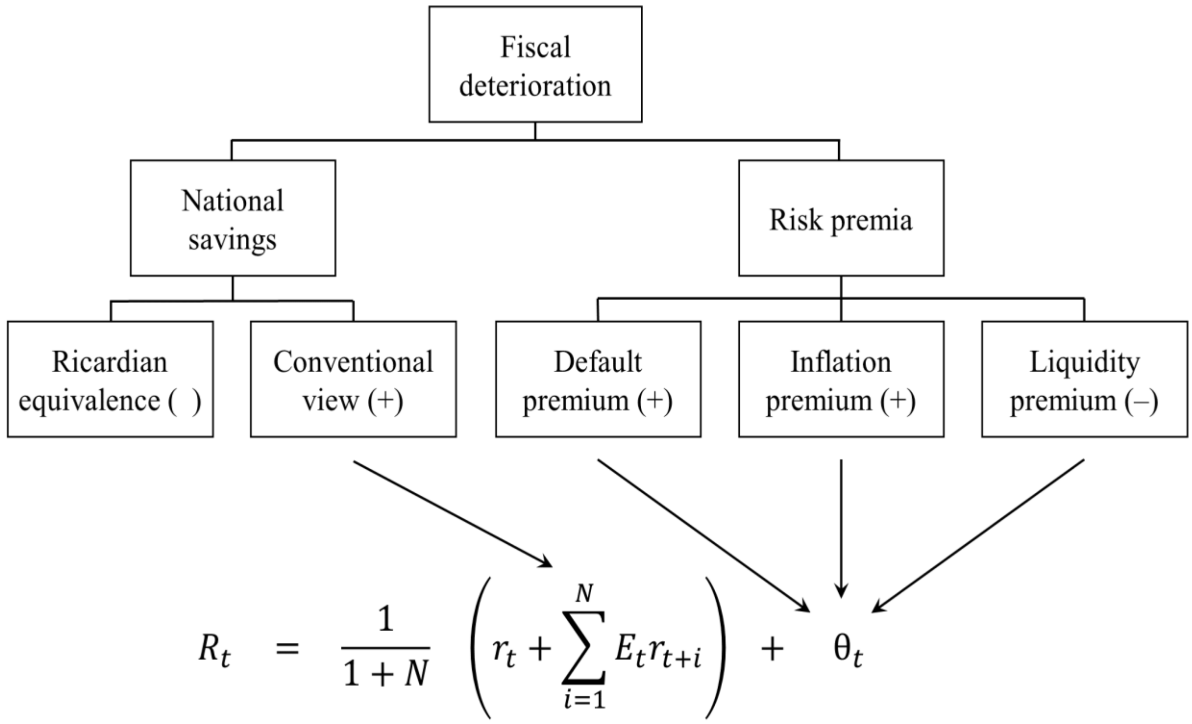


Figure 3. Expected interest rate responses to deterioration in fiscal balances (compiled by the author).

As was pointed out, the impact from fiscal deterioration on long-term interest rates can be expected to occur via a fall in national savings and a raise in risk premiums. The fall in savings, which ultimately results in a smaller capital stock yielding higher interest rate, raises the current and expected interest rates of all tenors and so affects the first sub factor of the above equation as a whole. In a closed economy, the magnitude of this effect is determined by how strongly private sector agents increase their savings as a response to the budget deficit, which is the point of contention between the conventional view and the Ricardian equivalence proposition. The effect through risk premiums can occur via three different risk premiums, from which default and inflation premiums are expected to raise and liquidity premium to lower the risk premium factor θ_t.

The openness of the economy and the stance of monetary policy also have an effect on long-term interest rates, although the effect is not a direct consequence from deterioration in fiscal balances like the influence from the fall in national savings and the raise in risk premiums. In an open economy, the externally funded share of the budget deficit dampens the crowding out effect of the fall in national savings and so lowers the pickup in long-term interest rates. Therefore, the degree of openness of the economy in effect determines how strongly the fall in national savings affects domestic long-term interest rates under the conventional view of debt. The stance of monetary policy, on the other hand, affects at least one of the two sub factors of the above equation: while the main monetary policy tool (i.e. the chosen policy rate) anchors the current short-term interest rates r_t, the employed monetary policy stance could also influence either the expected future short-term interest rates $\sum_{i=1}^{N} E_t r_{t+i}$ (with e.g. forward guidance (Del Negro et al., 2012, 2)) or the risk premium factor θ_t (with e.g. quantitative easing (Krishnamurthy & Vissing-Jorgensen, 2011, 3)), depending on the actual monetary policy framework employed by the central bank.

3 Review of the empirical literature

3.1 Challenges in the empirical analysis

The empirical literature on the subject is extensive, but is also known for its inconclusive evidence on the impact of public debt and deficits on interest rates (see e.g. Engen & Hubbard (2005), Ardagna et al. (2007) or Dell'Erba & Sola (2016)). Roughly one hundred papers have been conducted to analyse this relationship for the advanced economies, and yet it is still debated over whether the relationship even exists, not to mention its magnitude or under which conditions it prevails (Aisen & Hauner, 2013, 2501). There are a few major complicating factors that problematize the empirical analysis of this relationship. We note four of these in turn.

First, interest rates on government bonds are determined in financial markets where investors try to take all the relevant information that might affect their expected returns into account when deciding on whom to lend and at what rates. It is a difficult task to include all the necessary information which has affected the prevailing yields at the time to isolate the effect of fiscal policy on interest rates. Furthermore, due to the forward-looking nature of financial markets, interest rates rather depend on the expectations of fiscal policy and other variables than on their realized values (Gale & Orszag, 2004, 148). Gale and Orszag (2004) summarized the findings of 66 empirical studies and showed that papers which employ expected fiscal policy measures tend to find positive and statistically significant results more often than the ones that employ current measures. Out of the studies they covered, 47 % employing current measures found predominantly positive and significant effect while this figure was 53 % for the studies employing expected measures instead. Once studies that utilize VAR-based dynamics[15] to tackle the expectations were ruled out, the figure rose to 68 %. (Gale & Orszag, 2004, 186) While this finding is important and plausible, it brings about even more challenges to carrying out the empirical analysis as the market expectations about future fiscal policy measures are not directly observable but must be proxied instead. Proxies are again only reflections of expectations and may cause the coefficients on fiscal policy measures to be

[15] VAR-based projections are based on past values of the projected variables and so do not contain any other information but what is already embodied in the past values. For this reason, they are inferior to projections produced by government offices or intergovernmental organizations. (Gale & Orszag, 2004, 152)

biased towards zero due to measurement errors and so increase the tendency to underestimate the effects of fiscal policy on interest rates. (Gale & Orszag, 2004, 149)

A second complicating factor is that economic theories differ in their views regarding whether it is the flow (the budget balance) or the stock (the level of public debt) variable which should ultimately matter for the level of interest rates (Faini, 2006, 449). By and large, growth theories suggest that stock of public debt is the determining factor while business cycle theories underline the importance of budget balances. As a result, empirical studies have largely differed in which fiscal policy measure they choose to employ to test for the implications of public debt on interest rates. This largely complicates the comparison of the empirical findings. It is typical that the magnitude of the effect is found to be larger for budget deficits than for the stock of public debt. (Faini, 2006, 450).

A third complicating factor stems from business cycles, which induce reverse-causality effects between deficits and interest rates through countercyclical monetary policy and automatic fiscal stabilizers (Laubach, 2009, 859). During booms, automatic fiscal stabilizers decrease deficits by increasing tax revenue while interest rates simultaneously tend to rise due to the increased investment demand. As noted in section 2, the central bank can also be expected to be responsive to booms by raising policy rates to hold back the positive output gap and inflation, which further increases real interest rates in the economy. The reverse can be expected to happen during recessions. These mechanisms imply that both fiscal policy measures and interest rates are partly jointly determined by business cycles. This entails two further issues. First, an endogeneity problem is likely to be present as both independent and dependent variables of the model are possibly affected by an external factor, i.e. the business cycle. Second, this may cause the effect of deficits on interest rates to be masked behind seemingly negative correlation between deficits and interest rates, if the latter is larger in its magnitude than the former. To overcome these issues, empirical studies typically control their regressions for cyclical conditions e.g. by including control variables such as short-term interest rate, inflation rate, growth rate of GDP or output gap (e.g. Ardagna et al. (2007), Baldacci & Kumar (2010) & Chinn & Frankel (2007)), by using cyclically adjusted variables instead of actual ones (e.g. Ardagna (2009) & Faini (2006)), by averaging the data over longer periods (e.g. Aisen & Hauner (2013)), or by employing forward rates instead of current rates as the dependent variable, which can be expected to be less affected by cyclical factors (e.g. Laubach (2009) & Thomas & Wu (2009)).

The fourth and final complicating factor we note here is that there is no natural benchmark for the size of the interest rate effect of fiscal policy, like there is for the effect of taxes on consumption. This is because the size of the overall effect depends on various elasticities of different channels which might be expected to take domestic interest rates to different directions and so partly offset each other. For example, if openness of the economy or liquidity premium had a substantial impact on the domestic financial market, the overall effect of fiscal policy on interest rates would not need to be large or even exist, although the conventional view of debt would be valid. (Elmendorf & Mankiw, 1999, 1657–58) This makes the interpretation of the overall effect vague, especially if the magnitude of the effect is found to be small.

3.2 Findings of recent studies

The twelve recent empirical papers[16] that were examined for this subsection differ in their findings just as one would expect based on the statements that have been previously made in the literature to claim its inconclusiveness (see e.g. Aisen and Hauner (2013), Ardagna et al. (2007), Engen and Hubbard (2005) and Dell'Erba and Sola (2016)). After a closer look, this is not so surprising though; in addition to different results, the examined papers largely differ by their samples (by both the selection of countries and periods of time), methodological decisions (i.e. the model specification), employed variables (both dependent and independent) and so on.

Table 1 makes an effort to summarize the quantitative results of the examined papers. Most of them contain several estimation strategies whose results cannot be fit into one table. Thus, only the results from the baseline estimations with full samples are reported in table 1, as they are most suitable for cross-study comparison due to similar specifications. To highlight the key differences between the examined empirical papers, also their employed samples, main estimation methods as well as the independent variables are listed in table 1. Table 1 will be further utilized in section 5 when the empirical results of the thesis are compared to those of other studies.

[16] The papers that were selected for the cross-study comparison are all published after the summarizing study by Gale and Orszag (2004) so that our summary does not overlap with theirs. Additional benefit from utilizing only recent papers is that the time spans of those match much better with that of ours, which makes the results better comparable with our own empirical assessment.

Table 1. Summary of the findings in the recent empirical literature.

			Estimated impact on interest rates			
Paper (Year)	**Countries (Period)**	**Main method**	***Budget balance***	***Primary balance***	***Public debt***	***Change in public debt***
Aisen & Hauner (2013)	60 ADV & EME (1970–2006)	GMM	26			
Ardagna (2009)	16 OECD (1960–2002)	OLS		-124/162 $-97^{r}/142^{r}$		
Ardagna et al. (2007)	16 OECD (1960–2002)	DGLS		10 12^{r}	1 -3^{r}	
Baldacci & Kumar (2010)	31 ADV & EME (1980–2007)	FE	17 30^{r}	13 15^{r}		5 20^{r}
Chinn & Frankel (2007)	4 EMU, UK & US (1988–2006)	FE			6^{r}	$11^{r,e}$
Clayes et al. (2012)	35 OECD & EME (1990–2005)	LM			1	
Dell'Erba & Sola (2016)	17 OECD (1989–2013)	FAP		1^{e}	1^{e}	
Faini (2006)	9 EMU (1979–2002)	3SLS		3^{r}	0^{r}	
Gruber & Kamin (2012)	19 OECD (1988–2007)	FE		4^{e} $7^{r,e}$	1^{e} $1^{r,e}$	
Haugh et al. (2009)	10 EMU (2005–2009)	2SLS	$5^{e,s}$		0^{s}	
Laubach (2009)	US (1976–2006)	OLS	20^{e}	29^{e}	2^{e}	
Thomas & Wu (2009)	US (1983–2005)	OLS	$30–66^{r,e}$			

Notes: (1) Superscripts r and s denote real interest rates and interest rate spreads, respectively, while superscript e marks studies that employ expected fiscal measures instead of realized series. (2) All numbers are in basis points. (3) Numbers reported by Ardagna (2009) are cumulative effects over a five-year period and are divided into negative and positive effects (as responses to fiscal contractions and expansions).

Source: Compiled by the author.

The interest rate effect of the flow variables (budget balance, primary balance and change in public debt) reported in the papers varies immensely between 1 and 66 basis points while the equivalent of the stock variable (public debt) is found to be significantly smaller (and partly perverse), between -3 and 6 basis points. Studies utilizing both real and nominal rates as the dependent variable have found from 2 to 15 basis points higher impact from the flow variables when the real rate is employed. For the stock variable, the evidence is quite the contrary, 0 and -4 basis points. Interest rate responses to changes in primary balances have been found to be the lowest among the flow variables, which is reasonable given that they exclude interest payments and so are less subject to reverse causality than the other two measures. The quantitative results do not diverge with respect to the usage of expected fiscal measures versus realized values like one would expect based on the paper by Gale & Orszag (2004); in fact, once one excludes the studies based only on US data (Laubach (2009) and Thomas & Wu (2009)), it seems that the papers employing realized values have found quantitatively larger interest rate effects from fiscal policy.

In addition to providing the quantitative results of the baseline estimations, we pay attention to two other features that have been repeatedly noted in the literature and provide useful information for our own empirical workout. These are the nonlinearities of the relationship between fiscal policy and interest rates and the spill-over effects from global fiscal policy measures. There are, obviously, other relevant features as well (see e.g. Baldacci and Kumar (2010) on the influence of quality of institutions and country characteristics on this relationship) but as those do not fall under the scope of this thesis they will not be covered.

Many papers find the relationship between public debt (and to some extent, budget balances) and long-term interest rates to be nonlinear. For example, Baldacci and Kumar (2010) find that both high initial deficits (2 percent of GDP prior to the year of fiscal deterioration) and public debt-to-GDP ratios (above 60 percent) boost the impact of further fiscal deterioration by 14 and 6 basis points, respectively. Ardagna et al. (2007) arrive to somewhat similar conclusion by studying the squared terms of fiscal flow and stock variables as well as the interaction terms between their binary and continuous forms. While the magnitude of their findings is larger for primary balances, it is more robust for public debt-to-GDP ratios. Dell'Erba and Sola (2016) study the nonlinearities too but obtain far less clear evidence on them; they find the coefficient on public debt-to-GDP ratio to be different only for countries with 75 percent or higher public debt-to-GDP ratios, and even then, the impact is very small, only 0.3 basis points.

The economic reasoning behind the nonlinear relationship is patent, especially with respect to the level of public debt: high levels of public debt could raise concerns among investors about the solvency of the governments and so increase the default premium embodied in their bond yields, as discussed in subsection 2.2. Exceptionally low levels of public debt could increase the bond yields too if the outstanding amount of publicly traded government debt securities, i.e. the level of public debt, was not high enough to sustain a liquid government bond market. Rapid deterioration of budget balances could increase the default premium too, while it is not plausible to think that there would be a direct link between the liquidity premium and the budget balances.

Global fiscal policy measures are often found to have a larger effect on long-term interest rates than their domestic counterparts. Ardagna et al. (2007) use two different specifications to approximate for global measures and find that a one percentage point increase in global primary deficit and global public debt-to-GDP ratio increases long-term interest rates by 28–66 and 3–21 basis points, respectively. The inclusion of the global measures does not dispel nor significantly reduce the effect of the domestic measures in their study, implying that long-term interest rates are determined by both domestic and global fiscal policy measures. By concentrating only on EMU countries, Faini (2006) finds very similar results with the only noteworthy exception that the public debt-to-GDP ratio matters only at the EMU level (which corresponds to the "global measure" by its idea in their specification). There is also some opposing evidence about the effect from global measures; Gruber and Kamin (2012) find that foreign fiscal policy (sample total minus domestic) does not have the expected influence on interest rates. Instead, the sign varies across different specifications, making the overall evidence rather mixed.

The hypothesis that global fiscal policy affects domestic interest rates strongly relies on the idea that different government bonds are good substitutes to each other and that financial markets are deeply integrated. In extreme, i.e. when government bonds are perfect substitutes[17] to each other and when capital is perfectly mobile, this would mean that the nominal interest rate spreads between different government bonds would be fully reflected in exchange rate differentials so that the real interest rates would be equal across all countries. In this type of framework, domestic fiscal policies would not have any direct impact on domestic interest rates but would only influence them indirectly by affecting the pool of global savings which is then allocated to different government bonds in the above manner. (Ardagna et al., 2007, 10) Although the suggestion drawn from the extreme scenario is implausible due to a number of

[17] In our case, the perfect substitutability would mean that different government bonds were all equally risky assets and so did not differ by e.g. their liquidity characteristics or default probabilities.

reasons, the general idea is persistent: the better substitutes the government bonds are to each other and the more mobile the capital is, the larger the impact we may expect from global fiscal measures and the lesser the impact we may expect from crowding out of domestic capital.

By studying more in detail how the integration of financial markets influences the crowding out effect of public debt, Clayes et al. (2012) are able to show that the spill-over effects from global fiscal policies are rather limited in their magnitude. However, they become much stronger once the sample countries are narrowed down to specific country groups which can be expected to have eminently deep economic and financial integration. For example, among OECD and EU countries the spill-over effects are particularly strong. As a result, the authors argue that instead of the global spill-over effect it is in fact the cross-border spill-over effect that matters for the formation of domestic interest rates. While their result is plausible from a theoretical standpoint, to the knowledge of the author they are the only one to use such a specification and so unfortunately are not subject to any cross-study comparison.

4 Research data and methodology

4.1 Data and descriptive statistics

The thesis uses semi-annual panel data on 29 OECD economies covering a time span from the first half of 1989 until the first half of 2016. The goal was to include as many countries for as many years as possible to make the sample less vulnerable to selection bias. However, the decision to use OECD's Economic Outlook (EO) as the main data source narrowed down the list of sample countries to the following 29 OECD members: Australia, Austria, Belgium, Canada, Czech Republic, Denmark, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Italy, Japan, Korea, Luxembourg, Netherlands, New Zealand, Norway, Poland, Portugal, Slovak Republic, Slovenia, Spain, Sweden, Switzerland, the United Kingdom, and the United States of America.

The objective is to study the impact of fiscal policy on long-term interest rates, which for the following set of variables is utilized: realized nominal interest rate on 10-year government bonds (i^L), one-year-ahead nominal short-term interest rate (i^S), one-year-ahead inflation rate (π), one-year-ahead growth rate of real GDP (g), one-year-ahead primary balance-to-GDP ratio (PB), one-year-ahead public debt-to-GDP ratio (PD), realized central bank policy rate (CBR), realized sovereign credit rating (CR), government bond market size (GBM) as well as two alternative measures for capital account openness, ($KAOPEN$) and ($FAFLGDP$). The time series for realized nominal long-term interest rates are from Macrobond, whereas the fiscal and macroeconomic data are collected separately from OECD EO issues n. 45–99, which were published between June 1989 and December 2016 in semi-annual frequency. Data on central bank policy rates are from either Bloomberg or central bank websites, credit ratings are from Trading Economics and government bond market size values are calculated based on data from the 2017 June update of the database of Beck et al. (2000) and IMF World Economic Outlook Database, April 2017. The first measure of capital account openness, ($KAOPEN$), is from the 2017 July update of the database of Chinn and Ito (2006), while the second one, ($FAFLGDP$), is calculated based on data from the updated and extended version of the dataset constructed by Lane and Milesi-Ferretti (2007), and is further complemented with data from the International Financial Statistics (IFS) database of IMF. Table 2 provides the description of the variables as well as reports the data sources.

Table 2. Description of variables and data sources.

Variable	Description	Source
$\boldsymbol{i^L}$	Realized nominal interest rate on 10-year government bonds, monthly average yield.	Macrobond
$\boldsymbol{i^S}$	One-year-ahead nominal short-term (three-month) interest rate.	OECD Economic Outlooks n. 45–99
$\boldsymbol{\pi}$	One-year-ahead inflation rate, calculated based on GDP deflator (PGDP) values: $\ln(\frac{\mathrm{PGDP_{t+1}}}{\mathrm{PGDP_t}})$.	OECD Economic Outlooks n. 45–99
$\boldsymbol{g}$	One-year-ahead growth rate of real GDP, calculated based on real GDP (GDPV) values: $\ln(\frac{\mathrm{GDPV_{t+1}}}{\mathrm{GDPV_t}})$.	OECD Economic Outlooks n. 45–99
$\boldsymbol{PB}$	One-year-ahead primary balance-to-GDP ratio.	OECD Economic Outlooks n. 45–99
$\boldsymbol{PD}$	One-year-ahead public debt-to-GDP ratio.	OECD Economic Outlooks n. 45–99
$\boldsymbol{CBR}$	Realized nominal central bank policy rate.	Bloomberg & central bank websites
$\boldsymbol{CR}$	The average of the numeric conversions of credit ratings by Moody's, S&P and Fitch: from 0 (lowest) to 16 (highest).	Trading Economics
$\boldsymbol{GBM}$	Size of the government bond market in billions of dollars, calculated based on public bond market capitalization to GDP (pubond) values: $\mathrm{pubond} * \mathrm{GDP(US\$)}$.	Beck et al. (2017) & IMF WEO (April 2017)
$\boldsymbol{KAOPEN}$	An index measuring a country's degree of capital account openness; from 0 (lowest) to 1 (highest).	Chinn & Ito (2017)
$\boldsymbol{FAFLGDP}$	The sum of foreign assets (FA) and foreign liabilities (FL) as a share of GDP: $\frac{\mathrm{FA+FL}}{\mathrm{GDP}}$.	Lane & Milesi-Ferretti (2007) & IMF IFS

Source: Compiled by the author.

Realized nominal interest rate on 10-year government bonds is chosen as the dependent variable for two reasons. First, as noted in the subsection 2.3, long-term interest rates are likely to be less anchored by the chosen policy rate of the central bank than their short-term counterparts and so can be expected to better capture the possible effect from fiscal policy. Second, the OECD members have been issuing this type of debt security for many decades (Ardagna et al., 2007, 4), which makes it as minimally restrictive dependent variable as possible. The variable is observed as monthly averages of 10-year government benchmark bond yields.

Both one-year-ahead primary balance-to-GDP and one-year-ahead public debt-to-GDP ratios are used to measure the stance of fiscal policy, and so are the main independent variables of interest. Primary balance, which equals government net lending excluding interest payments on existing stock of government debt (OECD, 2005), is used instead of total budget balance in order to tackle reverse causality[18]. Gross public debt equals the cumulative amount of past budget balances and so provides an alternative measurement for fiscal policy with a longer-term perspective. Both fiscal policy indicators are one-year-ahead projections made by the OECD staff.

Short-term interest rates, inflation rates and growth rates of real GDP are all included as control variables; the first two are meant to account for the differences in the stances of monetary policy while the last one controls for business cycles. As discussed in 3.1, there are various ways of tackling the cyclical conditions, from among which including this type of set of control variables seems to be the most common approach (e.g. Ardagna et al. (2007) and Dell'Erba and Sola (2016) employ an identical set of control variables). Just like the fiscal policy indicators, the three control variables are one-year-ahead projections by the OECD staff. Expected growth rates of real GDP and inflation rates are not provided by the OECD directly but are instead calculated based on real GDP and GDP deflator values, respectively.

The remainder of the independent variables are all realized series instead of one-year-ahead forecasts. Central bank policy rates work as an alternative measurement for the prevailing stance of monetary policy. For central banks which employ policy rate ranges, e.g. the Federal Reserve System from December 2008 onwards, the realized nominal central bank policy rate is set to equal the average of the range. Credit ratings measure the sovereign risk. They are set to equal the average of the sovereign credit ratings assigned by the three big agencies, namely Moody's, S&P and Fitch, and are converted into numeric format in the following manner: B3/B-/B- equals 1, B2/B/B equals 2, and so on through Aaa/AAA/AAA equals 16[19]. This is a common method to transform bond ratings into data for regression analysis (Cantor & Packer, 1996, 41). Government bond market sizes are meant to proxy for the differences between the market liquidity of different government bonds. Public bond market capitalization-to-GDP ratios are multiplied by the dollar equivalents of the GDP values as to calculate government bond markets sizes that are comparable across different economies. Both measures for capital account openness tackle the differences between the degrees of financial openness across the

---

[18] Total budget balance includes interest payments, which are affected by the movement of interest rates. It is desirable to remove the effect of interest rates on budget balance when the opposite causal effect is being studied. Utilizing primary balances is an effortless way to address the issue.
[19] Appendix A presents the credit rating conversion methodology in more detail.

sample countries. While ($KAOPEN$) is a ready-made index by Chinn and Ito (2006), ($FAFLGDP$) is constructed as the sum of foreign assets and liabilities over GDP, which is proposed to be the *de facto* measurement for financial openness by Lane and Milesi-Ferretti (2003, 2007). The observations on central bank policy rates, government bond market sizes and the sums of foreign assets and liabilities over GDP are all complemented to reach the maximum data availability[20]. Furthermore, due to the unavailability of semi-annual time series for government bond market sizes and for both measures for capital account openness, the annual time series are used instead and are transformed into a semi-annual format so that the consecutive observations have the same values[21]. Table 3 provides the descriptive statistics.

Table 3. Summary statistics.

	N. of						
Variable	**obs.**	**countries**	**periods**	**Mean**	**S. D.**	**Min**	**Max**
$\boldsymbol{i^L}$	1281	17–29	12–55	5.24	3.17	-0.56	33.97
$\boldsymbol{i^S}$	1281	17–29	12–55	4.06	3.46	-0.79	24.50
$\boldsymbol{\pi}$	1281	17–29	12–55	2.17	1.49	-4.93	14.17
$\boldsymbol{g}$	1281	17–29	12–55	2.35	1.41	-9.80	7.73
$\boldsymbol{PB}$	1281	17–29	12–55	0.47	3.23	-11.91	14.44
$\boldsymbol{PD}$	1281	17–29	12–55	71.58	37.46	1.50	234.25
$\boldsymbol{CBR}$	1281	17–29	12–55	3.87	3.51	-0.75	24.50
$\boldsymbol{CR}$	1281	17–29	12–55	14.11	2.84	0.00	16.00
$\boldsymbol{GBM}$	1266	17–29	12–55	662.95	1701.56	0.00	12813.70
$\boldsymbol{KAOPEN}$	1195	17–29	12–55	0.91	0.18	0.17	1.00
$\boldsymbol{FAFLGDP}$	1281	17–29	12–55	10.14	39.02	0.63	374.75

Source: Author's calculations.

As Gale and Orszag (2004) showed, taking into account the forward-looking nature of financial markets is important and can in fact even be crucial in determining the outcome of the study in this field of research. Hence, we employ the one-year-ahead projections forecasted by the OECD staff for all the variables that they are available, i.e. for (i^S), (π), (g), (PB) and (PD).

[20] For central bank policy rates, the missing observations are substituted by one-year-ahead short-term interest rates, whereas the time-series of the government bond market sizes and the sums of foreign assets and liabilities over GDP are extended by proxying the missing values with the help of the public debt-to-GDP ratios and by utilizing data from the IFS database of IMF, respectively.

[21] Since the data from the annual series of (GBM), ($KAOPEN$) or ($FAFLGDP$) are only used to construct dummy variables, we are not concerned that the transformation from annual to semi-annual series biases our results.

For these variables, the data are collected so that from every single issue of OECD EO the one-year-ahead projections are stored to represent the expectations that prevailed at the time. For example, the forecasts for short-term interest rates from EO issue n. 45, released at June 1989, represent the market expectations for three-month money market rates for 1990 that prevailed during the first half of 1989. Furthermore, we follow the methodology used by Dell'Erba and Sola (2016) in the panel construction so that the average yields of 10-year government benchmark bonds are observed during the initial month after the releases of the forecasts; in July and in January for June and December releases, respectively. Figure 4 illustrates the panel construction.

Forecast		Forecasted period		Observation
$x'_{i,1989:\text{Jun}}$	$\rightarrow$	$x'_{i,1990}$	$\Longrightarrow$	$y_{i,1989:\text{Jul}}$
$x'_{i,1989:\text{Dec}}$	$\rightarrow$	$x'_{i,1990}$	$\Longrightarrow$	$y_{i,1990:\text{Jan}}$
$\vdots$	$\vdots$	$\vdots$	$\vdots$	$\vdots$
$x'_{i,2016:\text{Jun}}$	$\rightarrow$	$x'_{i,2017}$	$\Longrightarrow$	$y_{i,2016:\text{Jul}}$

Figure 4. Panel construction (compiled by the author).

In figure 4, x'_{it} denotes the set of independent variables, y_{it} is the dependent variable, i denotes the cross-sectional units, i.e. countries, and t stands for the time dimension, i.e. semi-annual periods. The advantage of this type of panel structure is that it reduces reverse causality as the realized average yields observed in the initial months after the releases of the forecasts cannot have an effect on the forecasts anymore. As a comparison, if one used e.g. average yearly yields instead, one would possibly encounter reverse causality as the prevailing market conditions were likely embodied in the forecasts that were made by the OECD staff (Dell'Erba & Sola, 2016, 400). For example, the June 1989 forecasts for 1990 were likely affected by the realized long-term interest rates, among other market conditions, that prevailed during the first half of 1989. As the yields observed in the first half of 1989 would account for roughly half of the average yearly yields, 50 percent of the information that was used to generate the observations on the dependent variables were in fact also part of the process that generated the independent variables. The main reason for the above panel structure is to overcome this issue.

The expectations-based time series are not available for the remainder of the variables, i.e. for (CBR), (CR), (GBM), $(KAOPEN)$ and $(FAFLGDP)$, which for realized series are used instead so that the observations match with those on the dependent variable. For central bank policy rates and sovereign credit ratings, forecasts could provide more valuable information than the realized series, but for government bond market sizes, as well as for both measures of capital account openness, this is less of an issue; the last three variables depict characteristics of an economy that develop rather slowly and are unlikely to be a target for the speculative behaviour of financial markets.

The final thing we note before moving forward is that the panel is unbalanced, i.e. that the number of observations in the sample is not i times t. The reason for this is simply the data availability; for 17 countries, the data are available for all the 55 semi-annual periods, whereas for the rest 12 countries only shorter time-series are available. Apart from being shorter, the time-series are complete, meaning that they do not have any missing observations within the series. The starting point of each series is set to whenever all the needed variables are in place. In general, fixed effects (FE) estimation[22] is consistent and asymptotically normal on unbalanced panels as far as the idiosyncratic errors u_{it} are mean independent of the selection indicators which determine the individuals that are selected in the sample (Wooldridge, 2010, 830). Hence, we are confident that the unbalance of the panel does not bias our results.

4.2 Projections versus actual series

We look at the employed projections slightly more in detail as they play such a crucial role in the panel construction. The type of dataset that utilizes information that was available to decision makers (in our case, to investors) at the time of decision-making (investment decisions) is often referred to as a "real-time dataset" in the literature (see e.g. Beetsma and Giuliodori (2010), Cimadomo (2012) and Dell'Erba and Sola (2016)). It has the obvious advantage over ex post data as it can be expected to better match the information that the decisions were based on (Beetsma & Giuliodori, 2010, 420).

However, utilizing real-time data is by no means trouble-free, especially whenever it is applied to approximate for fiscal expectations. Foremost, this is because fiscal projections from national authorities may be subject to political bias, meaning that policy-makers may have an incentive to publish fiscal plans that are different from the "true" fiscal path that is most likely going to be realized in the implementation stage of fiscal policy (Beetsma & Giuliodori, 2010,

[22] FE-estimation is covered more in detail when the model selection is carried out in subsection 4.3.

421). For example, fiscal and growth projections reported in the Stability and Convergence Programmes (SCPs)[23] by EU governments operating under delegation are found to be systematically upward-biased while the reverse is found to hold for governments operating under strong fiscal rules (von Hagen, 2010, 501).

Employing projections from an independent institution like OECD can be expected to decrease the political bias as the OECD staff is not likely to share the same incentives with national authorities. In fact, the mean errors by OECD for projecting budget deficits are usually found to be below 0.5 points, even though "balance variables" (e.g. current accounts or budget balances) are among the least well forecasted macroeconomic variables. Even more importantly, systematic under- and over-predictions of budget deficits by OECD are found to be rare. (Artis & Marcellino, 2001, 34) For this reason, we do not have a major concern that the employed projections would be unsuitable to account for market expectations that prevailed at the time. In addition to being less politically biased, there are at least two other advantages using OECD data over data from national authorities. First, the methodological decisions for constructing the variables are the same for all countries, which makes the observations better comparable. Second, the data from OECD also enables a larger panel in both cross-section and time-series dimensions than data from SCPs. (Beetsma & Giuliodori, 2010, 422)

Next, we compare all the employed projections to their revised values from OECD EO n. 100 to see how the forecasts have performed against the actual series. This rough comparison is done by plotting the medians of the June and December projections against the equivalents from the revised data as well as by looking at the descriptive statistics of the forecast errors, which are simply the differences between the forecasts and the revised values. This type of comparison allows for the assessment of June and December projections separately and so also helps us to evaluate whether it is reasonable to pool the different issues together. The sample size was slightly cut (in total,71 observations were dropped) so that all the three subsamples would have the same amount of observations and so would be identically unbalanced[24]. Figure 5 (panels A to E) plots the projections against the realized series while table 4 reports the descriptive statistics of the forecast errors.

[23] SCPs represent an example of another concentrated database for our needs. Its data is based on reports by national authorities to European Comission (EC) and is used to evaluate how well EU Member States perform against their Medium-Term Budgetary Objectives (MTOs) (European Commission).

[24] If the three subsamples were not identically unbalanced, a different number of countries would be used to calculate the medians for some periods, which could bias the comparison based on the medians.

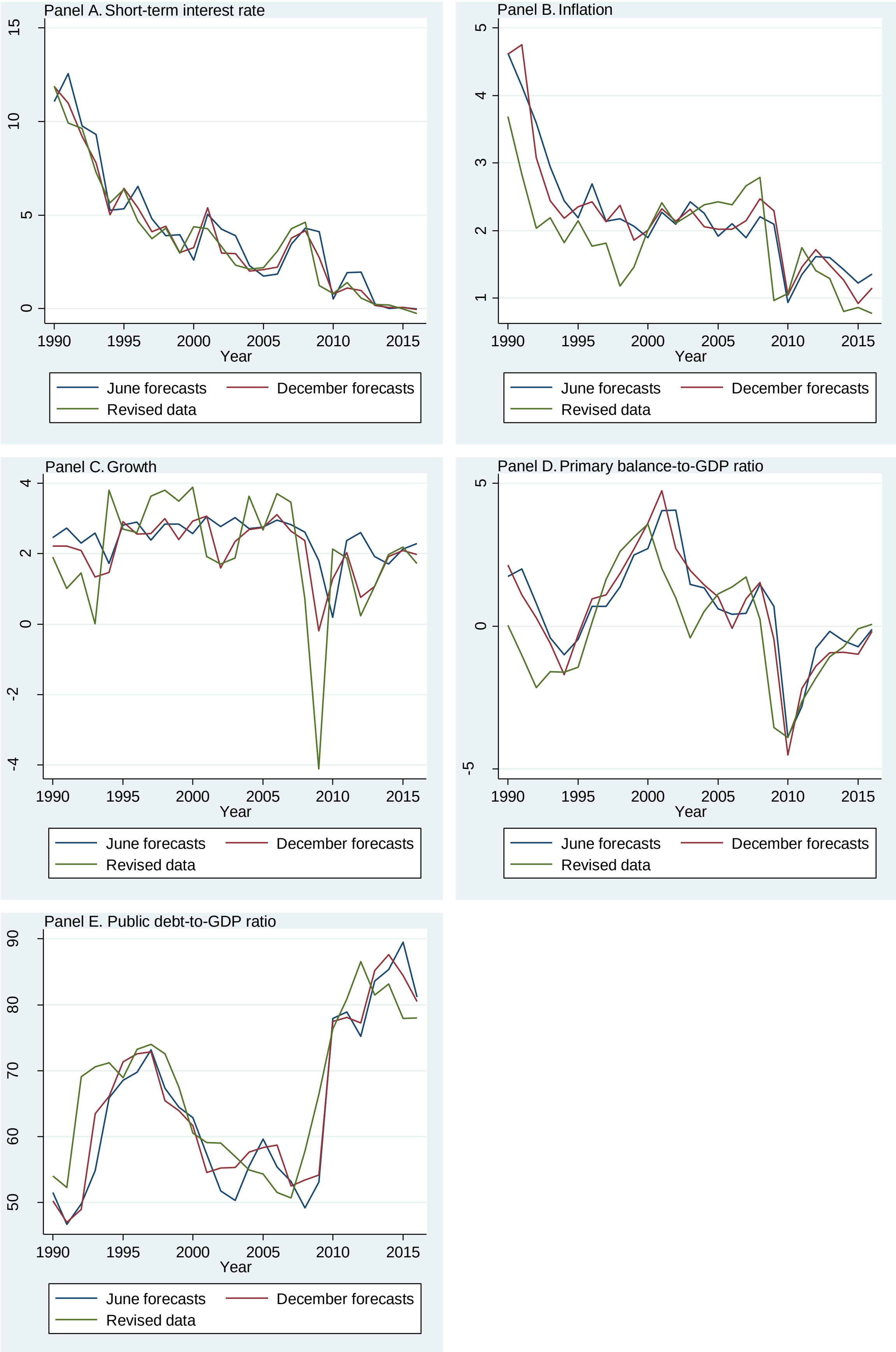


Figure 5. Forecasts vs. actual values (author's calculations based on data from OECD Economic Outlooks.

Table 4. Summary statistics of forecast errors.

Variable	Issue	Mean	S. D.	Min	Max
Short-term interest rate	June	-0.39	1.35	-5.59	5.34
	December	-0.15	0.89	-5.81	2.86
Inflation	June	-0.10	1.57	-7.46	11.94
	December	-0.11	1.38	-6.05	9.44
Growth	June	-0.52	2.41	-11.47	21.20
	December	-0.10	1.96	-9.45	20.09
Primary balance	June	-0.70	2.89	-17.90	16.20
	December	-0.55	2.66	-18.98	16.60
Public debt	June	3.54	10.85	-41.28	62.75
	December	2.61	10.04	-51.73	43.88

Source: Author's calculations.

What figure 5 (panels A to E) and table 4 imply does not come as a surprise: the 29 OECD members have on average run looser fiscal policies, experienced smaller growth and inflation rates and lower short-term interest rates than was forecasted by the OECD a year earlier. However, the forecasts have been relatively on target on average, consolidating further that OECD projections are suitable to account for market expectations. The differences between June and December projections seem reasonable as well. It is plausible to think that December projections would be more accurate than their June counterparts as the forecasted periods are half a year closer, which is indeed the case. The timing of December projections also makes it more likely (for some of the sample countries) that the OECD forecasts are based on official budget documents approved by national parliaments (Beetsma & Giuliodori, 2010, 422–423), which should further enhance the accuracy of the December projections.

The final thing to note is that the forecast errors have occasionally been quite extreme. For example, the maximum forecast error for public debt-to-GDP ratio is nearly 63 percent. However, once the country-period combinations which produced these minimum and maximum forecast errors are observed, a plausible explanation is quickly found: most of the extreme errors come from Iceland, Ireland and Slovak Republic and are related to either financial crisis or its aftermath. All three are among the countries which experienced outstandingly harsh economic development as a response to the crisis. As the magnitude of this development was in all likelihood not accurately foreseen, it is reasonable to think that this is how the forecast errors should look like for these country-period combinations.[25]

[25] Appendix B lists the country-period combinations which produce the minimum and maximum forecast errors for each variable.

4.3 Model specification

Panel data analysis is designed for longitudinal data sets which contain both cross-section and time-series dimensions, and so have the following form:

$$x_{it}, \qquad i = 1, \dots, N;\ t = 1, \dots, T,$$

where i is the cross-section dimension and t is the time dimension. Longitudinal data sets have several advantages over cross-section or time-series data sets, from which Hsiao (2003, 3–7) lists five. They:

1) *usually contain a large number of observations and so improve the efficiency of econometric estimates by providing higher degrees of freedom and reducing the collinearity among explanatory variables,*
2) *enable the analysis of many research questions that are otherwise not addressable,*
3) *provide a way to solve or reduce the omitted variable problem by utilizing both individual and time variation of data,*
4) *may simplify computation and interpretation of econometric estimates, and*
5) *enable more accurate predictions for individual outcomes than time-series data alone if the partial effect of on the dependent variable is similar across different individuals.*

The third property is particularly interesting for the thesis, as it is very likely that the omitted variable bias will be present in a simple multivariable reduced-form regression that is used to model the impact of fiscal policy on interest rates. However, if the employed model allows the unobserved variables[26] to be correlated with the observed ones, i.e. $Cov(a_i, x_{it}) \neq 0$, we might be able to solve, or at least reduce the omitted variable bias in the sample, and so obtain less biased results.

The presence of unobserved effects may be expressed in a simple unobserved effects model (Wooldridge, 2010, 285). For a panel of N cross-sectional units over T periods of time, the model has the form:

---

[26] A time-constant and unobserved variable is called an unobserved effect in panel data analysis (Wooldridge, 2010, 282).

$$y_{it} = \alpha_i + x'_{it}\beta + \varepsilon_{it}, \tag{1}$$

where y_{it} is the dependent variable, x'_{it} is a K-dimensional row vector of observed independent variables, β is a K-dimensional column vector of parameters, ε_{it} is an idiosyncratic error term and α_i is the intercept, which may vary across cross-sectional units. The model (1) may be expressed more conveniently with vectors and matrices:

$$y_i = e_T\alpha_i + X_i\beta + \varepsilon_i, \tag{2}$$

where y_i is a T-dimensional column vector of the y_{it}, e_T is an identity matrix of size T, α_i is a T-dimensional column vector of intercepts, X_i is a matrix of size $T{\times}K$ whose tth row is x'_{it} and ε_i is a T-dimensional column vector of errors. Writing (2) in a matrix form, we obtain:

$$\underset{(NT\times 1)}{\underset{y}{\begin{bmatrix} y_1 \\ y_2 \\ \vdots \\ y_N \end{bmatrix}}} = \underset{(NT\times N)}{\underset{D_N}{\begin{bmatrix} e_T & 0 & \cdots & 0 \\ 0 & e_T & \cdots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \cdots & e_T \end{bmatrix}}} \underset{(N\times 1)}{\underset{\alpha}{\begin{bmatrix} \alpha_1 \\ \alpha_2 \\ \vdots \\ \alpha_N \end{bmatrix}}} + \underset{(NT\times K)}{\underset{X}{\begin{bmatrix} X_1 \\ X_2 \\ \vdots \\ X_N \end{bmatrix}}}\beta + \underset{(NT\times 1)}{\underset{\varepsilon}{\begin{bmatrix} \varepsilon_1 \\ \varepsilon_2 \\ \vdots \\ \varepsilon_N \end{bmatrix}}}, \tag{3}$$

or more simply:

$$y = D_N\alpha + X\beta + \varepsilon. \tag{4}$$

(Mátyás, 2008, 24)

Intercept α_i is not observed but is instead estimated along the slope parameters β of the model. The omission of relevant independent variables, i.e. the presence of unobserved variables, may cause the intercept to be biased. For the same reason, the intercept actually contains any variation in the dependent variable y_{it} that cannot be explained by the set of observed independent variables x'_{it}. Because of this, the intercept α_i contains in essence the unobserved effect in the model.

Panel data analysis offers a variety of ways to deal with the unobserved effects. To tackle the issue, the thesis considers two possibilities, random effects (RE) and fixed effects (FE) estimations, which are among the simplest panel data models. The presence of unobserved effects is not taken for granted, and so the application of the pooled ordinary least squares

(POLS)[27] estimation is also considered. As a result, we have three candidate models, among which the selection ultimately depends on the quality of the intercept α_i, i.e. the unobserved effect[28]. The selection process between the candidate models can be simplified in the following manner:

1) *POLS, if* $\mathrm{\alpha_i = \alpha_0}$,
2) *RE, if* $\mathrm{\alpha_i = u_i; u_i \sim IID(\alpha_0, \sigma_u^2)}$, *and*
3) *FE, if* $\mathrm{Cov(\alpha_i, x_{it}) \neq 0}$.

If the intercept α_i is the same for all i, i.e. there is no unobserved heterogeneity among the cross-sectional units, then the pooled ordinary least squares estimation is applicable and produces consistent estimators of the slope parameters β (Wooldridge, 2010, 291). Practically speaking, this approach disregards the longitudinal nature of the data and simply treats the entire sample as a single cross section.

Random effects estimation allows for unobserved heterogeneity but requires zero correlation between the unobserved and the observed explanatory variables, i.e. $Cov(\alpha_i, x_{it}) = 0$ (Wooldridge, 2010, 286). To meet this requirement, it produces independent and identically distributed intercepts α_i, which have a constant expected value $E(\alpha_i | x'_{it}) = \alpha_0$ and variance $Var(\alpha_i | x'_{it}) = \sigma_u^2$, given all independent variables x'_{it}. Random effects estimation should be applied whenever its strict requirements are met, as it produces the most efficient estimators under these circumstances (Kennedy, 2003, 305–312).

Fixed effects estimation allows for unobserved heterogeneity so that independent variables x'_{it} may be arbitrarily correlated with the intercept α_i, i.e. $Cov(\alpha_i, x_{it}) \neq 0$. It is more robust than random effects estimation as it can estimate the slope parameters β consistenly even in the presence of time-constant omitted variables. The drawback of fixed effects estimation is that one is not able to include time-constant independent variables in the model since they cannot be distinguished from time-constant unobserved effects. (Wooldridge, 2010, 300–301)

[27] The usual ordinary least squares estimation is called pooled ordinary least squares estimation whenever it is applied on longitudinal data sets. (Wooldridge, 2010, 169–170).
[28] See e.g. Wooldridge (2012) for a full list of assumptions needed for RE and FE estimations.

In practice, the selection between the candidate models is done by running both F- and Hausman tests for the baseline regression[29], wherein the set of independent variables consists of the semi-annual projections for short-term interest rates (i^S), inflation (π) and growth (g) rates as well as primary balance-to-GDP (PB) and public debt-to-GDP (PD) ratios for $year_{t+1}$, and so the K-dimensional row vector of independent variables x'_{it} in (1) becomes

$$x'_{it} = [i^S{}_{it} \quad \pi_{it} \quad g_{it} \quad PB_{it} \quad PD_{it}]. \tag{5}$$

First, an F-test is applied to test the hypothesis that the intercept $\alpha_i = \alpha_0$ for all i against the alternative hypothesis that $\alpha_i \neq \alpha_0$ for at least one i. It compares the goodness-of-fit measures of a constrained estimation (here POLS) and an unconstrained estimation (here FE) to test whether there is any unobserved heterogeneity present in the model. It has the following form:

$$F(N-1, NT-N-K) = \frac{(SS_{POLS} - SS_{FE})/(N-1)}{SS_{FE}/(NT-N-K)}, \tag{6}$$

where the static has a distribution of an F-variable with $(N-1)$ and $(NT-N-K)$ degrees of freedom, SS_{FE} is the sum of squared residuals for the FE-estimator, and SS_{POLS} is the equivalent for the POLS-estimator. (Mátyás, 2008, 28–29) With the full sample size and for the explanatory variables defined above, the $F(28, 1247)$ has a value of 13.29, while the critical value at the 0.05 level of significance is 1.46. Hence, the H_0 is rejected, implying that there is a significant unobserved effect or at least a significant increase in the goodness-of-fit measure once the FE-estimator is applied over the POLS-estimator.

Second, to confirm that the presence of the unobserved effect should be tackled by using an FE-estimator instead of an RE-estimator, the Hausman test is performed. Essentially, it compares the slope parameters β of the FE- and the RE-estimators (in fact, of any two estimators being tested) under the null hypothesis that either of the two is consistent, while the alternative hypothesis is that only the FE-estimator is consistent. It has the form:

$$Q_H = (\hat{\beta}_{FE} - \hat{\beta}_{RE})' \left[\widehat{V(\hat{\beta}_{FE})} - \widehat{V(\hat{\beta}_{RE})}\right]^{-1} (\hat{\beta}_{FE} - \hat{\beta}_{RE}), \tag{7}$$

[29] This set of independent variables is widely-used in the literature and so forms the so-called baseline regression in our analysis. For this reason, the model selection is also performed under this specification.

where the static has a Chi-Squared distribution with dim(β) degrees of freedom, $\hat{\beta}_{FE}$ are the slope parameters of the FE-estimator and $\hat{\beta}_{RE}$ are the equivalent for the RE-estimator. (Mátyás, 2008, 81) With the same specification as above, the Q_H has a value of 38.07, which is larger than the fractile of the X_5^2 distribution at the 0.05 significance level, which equals 11.07. Thus, the H_0 is rejected, meaning that the slope parameters β of the FE-estimator are significantly different from the ones of the RE-estimator. This indicates that unobserved effects are correlated with the set of independent variables and so the FE-estimator is preferred over the RE-estimator. Based on the results of both F- and Hausman tests, we conclude that the FE-estimator better captures the variation in the long-term interest rates than either the $POLS$- or the RE-estimator.

Finally, we briefly note that employing an FE-estimator also enables the usage of a time varying intercept λ_t, which may be arbitrarily correlated with the set of observed independent variables x'_{it}, i.e. $Cov(\lambda_t, x'_{it}) \neq 0$. By adding λ_t to model (1), it becomes:

$$y_{it} = \alpha_i + \lambda_t + x'_{it}\beta + \varepsilon_{it}. \tag{8}$$

(Mátyás, 2008, 29) The implication of λ_t is similar to that of α_i: λ_t contains any variation in the dependent variable y_{it} that cannot be explained by x'_{it} and is common to all observations at given time t. The benefit from extending model (1) to also tackle the individual-constant unobserved variation in y_{it} is obvious as we aim at excluding all the variation that cannot be explained by x'_{it}. There are, however, at least two drawbacks in employing these so-called time fixed effects in an FE-estimation. First, we cannot identify the cause of individual-constant unobserved variation in y_{it} any more accurately than by linking it to a specific period of time t. Second, utilizing individual-constant independent variables together with time fixed effects is not an option as their effects cannot be distinguished from each other, as is the case with the usage of time-constant independent variables in an FE-estimation.

4.4 Time-series properties

As we are employing a long panel with 55 time-periods, at most, and 29 cross-sectional units, it is plausible to think that the data has either large-T asymptotics or both large-N and -T asymptotics but not large-N asymptotics alone, which is the case for so called micro panels.

This means that the panel basically resembles the usage of multiple time-series rather than that of multiple cross-sectional draws. The idea becomes clearer once the main channel of impact is being thought over again: the response of long-term interest rates to *e.g. loosening fiscal policy via national savings* is a time-series process where an economy gradually adjusts to a new equilibrium growth path with lower capital stock yielding a higher interest rate.

Due to this feature of the panel, it is important to examine the time-series properties of the main variables of interest, i.e. whether the variables can be treated as stationary or not, as these affect the statistical properties of time-series estimators: Under stationarity, the estimators' limiting distributions approximate to normal when T tends to infinity while under non-stationarity, or when data contain unit roots, the distributions will be nonstandard when T tends to infinity (Hsiao, 2003, 298). Even more importantly, we want to rule out the possibility of running spurious regressions, which would be the case if regressions were run between at least two $I(1)$ random vectors, e.g. y_i and one of x_{it}', without a co-integrating relation between them (Phillips & Moon, 1999, 1058).

We begin by examining the unit roots by applying both Im-Pesaran-Shin (2003) (IPS) and Fisher-type (with "dfuller" option) tests to the baseline regression[30], as these are, to the knowledge of the author, the only two unit root tests suitable for unbalanced panels in Stata 14.2. Both tests are based on the (augmented) Dickey-Fuller (ADF) regression

$$\Delta y_{it} = \phi_i y_{i,t-1} + z_{it}'\gamma_i + \epsilon_{it}, \tag{9}$$

where y_{it} is the variable being tested, $z_{it}'\gamma_i$ is the panel-specific mean and linear time trend (whenever decided to be included) and ϵ_{it} is a stationary error term, and are used to test the null hypothesis of unit root, i.e. $H_0\colon \phi_i = 0$ for all i, against the alternative hypothesis that $H_a\colon \phi_i < 0$ for a fraction of all i (in the case of IPS), or that $H_a\colon \phi_i < 0$ for at least one i (in the case of Fisher-type ADF test). (StataCorp., 2015, 516–526) (See StataCorp. (2015) for more detailed information on both tests.)

[30] We study the time-series properties of the data only for the variables included in the baseline regression as the rest of the variables, with the exception of credit rating, are only used to construct dummy variables which can be expected not to affect the time-series properties of the data. Credit rating, on the other hand, is a categorical variable and so not applicable to unit root testing.

First, both tests are applied to each variable with and without the linear time trend parameter and after this, both tests are applied again on the first differences of those variables that the null hypothesis of unit root could not be rejected with certainty. The lag structure for the IPS tests is chosen so that the Akaike information criterion (AIC) is minimized for each regression. The IPS test allows for a different lag structure for each time-series and reports the average number of lags for the entire panel in the output. For the Fisher test, the lag structure is chosen so that it equals the closest integer to the average number of lags of the respective IPS test. The upper half of table 5 reports the results of the unit root tests, while the lower half is devoted for the results of the Westerlund (2007) co-integration tests that are applied to the six-variable relation later.

Table 5. Results of unit root and cointegration tests.

	Im-Pesaran-Shin			**Fisher-type ADF**			
V.	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	
i^L	2.16	-3.74***	-33.19***	37.73	128.08***	723.92***	
	(1.07)	(1.10)	(1.52)	(1.00)	(1.00)	(2.00)	
i^S	-0.51	-4.04***		163.91***	167.21***		!
	(0.90)	(1.00)		(1.00)	(1.00)		
π	-10.00***	-10.33***		181.30***	138.30***		
	(0.83)	(0.83)		(1.00)	(1.00)		
g	-6.08***	-7.24***		193.45***	234.61***		
	(1.97)	(1.86)		(2.00)	(2.00)		
PB	-5.31***	-3.33***		99.85***	62.83		!
	(0.72)	(1.00)		(1.00)	(1.00)		
PD	2.09	0.55	-23.25***	32.24	35.57	455.91***	
	(0.90)	(0.86)	(0.79)	(1.00)	(1.00)	(1.00)	
Trend	No	Yes	No	No	Yes	No	
1st diff.	No	No	Yes	No	No	Yes	

Westerlund cointegration tests			
G_τ	G_α	P_τ	P_α
-3.45**	-24.63**	-18.62***	-22.22***
(1.00)	(1.00)	(1.00)	(1.00)

Notes: (1) Significance codes are the following: 0.1 '#', 0.05 '*', 0.01 '**' & 0.001 '***'. (2) Number of lags are reported in parenthesis. (3) Conflicting results between the two tests are marked with an exclamation mark. (4) All six variables are included in the four cointegration tests. (5) Hungary, Slovenia and Switzerland are excluded in the cointegration tests due to the shortage of data.

Source: Author's calculations.

The evidence is in favour of stationarity for short-term interest rates[31], inflation and growth rates and primary balances already without a deterministic trend. Once the trend is added, long-term interest rates also become stationary series. The trend stationarity for long-term interest rates is not a surprising find given that they declined in almost every country in the sample from the beginning of 1990s till the first half of 2016. While this is the case in our sample, it is not plausible to think of this applying to the population, as this would imply that long-term interest rates would be trend stationary processes with downward sloping deterministic trends at any given period of time. For this reason, the two unit root tests are also applied to the first difference of long-term interest rates and by doing so the variable is identified to be an $I(1)$ process. While the above evidence does not provide a strict guidance on whether we should treat long-term interest rates as a trend stationary process or as an $I(1)$ process, the economic rational would tilt more towards the latter option.

The only variable with strict evidence against stationarity is the public debt-to-GDP ratio, which is identified to be an $I(1)$ process in both unit root tests. Now, in case also long-term interest rates are an $I(1)$ process (out of the two alternatives above), there is a chance of running a spurious regression if there is no co-integrating relation between the panel variables. To be confirmed that this is not the case, we apply the four panel co-integration tests developed by Westerlund (2007) to the panel. All of the tests are based on the following data-generating process:

$$\Delta y_{it} = \delta_i' d_t + \alpha_i y_{i,t-1} + \lambda_i' x_{i,t-1} + \sum_{j=1}^{p_i} \alpha_{ij} \Delta y_{i,t-j} + \sum_{j=-q_i}^{p_i} \gamma_{ij} \Delta x_{i,t-j} + e_{it} \qquad (10)$$

where λ_i' equals $-\alpha_i \beta_i'$, d_t accounts for the deterministic components, p_i and q_i are the lag and lead orders, respectively, and α_i is the error correction parameter, which equals the speed of returning to equilibrium relation for the system. The null hypothesis to be tested is that there is no error correction, or co-integration, in the panel, i.e. $H_0\colon \alpha_i = 0$ for all i, against the alternative hypothesis that the panel variables are co-integrated, i.e. $H_a\colon \alpha_i < 0$ for at least one

[31] The IPS test for short-term interest rates was run in total with nine different lag structures and the one reported in table 5 is in fact the only one where the null hypothesis cannot be rejected at the 0.1 level. Additionally, the Fisher-type ADF test allows the rejection of the null hypothesis at the 0.001 level. Hence, we conclude that the evidence is more in favour of stationarity for short-term interest rates in levels.

i (in the case of G_τ and G_α), or that $H_a: \alpha_i < 0$ (in the case of P_τ and P_α).[32] (Persyn & Westerlund, 2008, 233–234) (See Westerlund (2007) and Persyn and Westerlund (2008) for more detailed information on the test statistics).

Westerlund (2007) cointegration tests are run for the six-variable relation (i.e. the baseline regression) so that a constant and a time trend are included and the lag order is set to one. This type of specification requires at least 22 observations per time-series and so the three shortest ones (Hungary, Slovenia and Switzerland) have to be excluded in order to make the test available. Increasing the lag order from one would increase the number of minimum observations per time-series and so would result in discarding even more countries from the sample. For this reason, we content ourselves with the lag order of one instead of selecting an AIC optimal lag order, as was done with the unit root tests. The results of the Westerlund (2007) co-integration tests all imply that there is a co-integrating relationship between the main variables of interest. Hence, we conclude that we can estimate the regression in levels without the danger of running a spurious regression.

[32] The difference between G and P test statistics are in their treatment of homogeneity of α_i; the group-mean tests (G_τ and G_α) do not require α_i to be homogenous across the panel while the panel tests (P_τ and P_α) do (Persyn & Westerlund, 2008, 233). The difference between τ and α test statistics are that the latter are normalized by the number of time periods t (Westerlund, 2007, 717).

5 Empirical results

5.1 Baseline regressions

The baseline regression consists of public debt-to-GDP (PD) and primary balance-to-GDP (PB) ratios, short-term interest rates (i^S), inflation (π) and growth (g) rates, and is used to explain the variation in long-term interest rates (i^L). All models are estimated in levels with an FE-estimator, wherein we always include country fixed effects but take turns employing time trend parameters and time fixed effects[33]. Table 6 reports the results of the baseline regressions.

Table 6. Long-term interest rates and fiscal policy – baseline regressions.

	Full sample 1989:1–2016:1		**First half 1989:1–2007:1**		**Second half 2007:2–2016:1**	
Variable	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
i^S	0.522	0.495	0.612	0.642	0.354	0.229
	(0.052)***	(0.058)***	(0.052)***	(0.057)***	(0.073)***	(0.074)**
π	-0.111	-0.116	0.276	0.245	-0.163	-0.195
	(0.100)	(0.109)	(0.102)*	(0.078)**	(0.134)	(0.134)
g	-0.353	-0.412	0.106	0.081	-0.332	-0.545
	(0.188)#	(0.227)#	(0.059)#	(0.055)	(0.186)#	(0.250)*
PB	-0.068	-0.012	-0.132	-0.082	0.054	0.065
	(0.032)*	(0.035)	(0.036)***	(0.024)**	(0.060)	(0.058)
PD	0.015	0.011	0.009	0.005	0.024	0.022
	(0.008)#	(0.006)#	(0.003)**	(0.002)*	(0.017)	(0.016)
Country FE	Yes	Yes	Yes	Yes	Yes	Yes
Time FE	No	Yes	No	Yes	No	Yes
Time trend	Yes	No	Yes	No	Yes	No
$\mathbf{R^2}$	0.741	0.792	0.901	0.950	0.389	0.449
N. of periods	12–55	12–55	3–37	3–37	12–18	12–18
N. of countries	29	29	28	28	29	29
N. of obs.	1281	1281	765	765	516	516

Notes: (1) Significance codes are the following: 0.1 '#', 0.05 '*', 0.01 '**' & 0.001 '***'. (2) Standard errors (in parenthesis) are corrected for heteroscedasticity by using vce(robust) option in Stata 14.2. (3) 'Country FE', 'Time FE' and 'Time trend' denote the usage of country fixed effects, time fixed effects and linear and quadratic time trends, respectively.

Source: Author's calculations.

[33] Due to the strong trending behaviour of long-term interest rates in our sample, it is reasonable to include at least linear and quadratic time trend parameters in the models. Switching from time trend parameters to time fixed effects allows us to observe whether our results are robust to time-specific global factors that are not directly studied in the thesis.

The results show that with the full sample, one percentage point deterioration in primary balances increase the long-term interest rates by nearly seven basis points, while the effect from an equivalent raise in public debt-to-GDP ratio is two basis points (column 1). Both coefficients are statistically significant, primary balances at the 0.05 level and public debt-to-GDP ratio at the 0.1 level. Compared to the findings of other studies (subsection 3.2, table 1), the response to primary deficits is from the low-end (other studies report coefficients ranging from 1 to 29 basis points), whereas the response to public debt-to-GDP ratio is in line with other studies.

There are a few disconcerting features in the results. The coefficient of primary balance-to-GDP ratio is not robust to switching from time trend parameters to time fixed effects (column 1 vs. 2). Instead, the time fixed effects crowd out the impact of primary balances on long-term interest rates; its magnitude decreases down to one basis point and it becomes statistically insignificant. The effect of public debt also slightly declines but maintains its significance. The fact that the impact of fiscal policy largely disappears under time fixed effects implies that there are time-specific factors that are omitted in the model and explain an important share of the variation in long-term interest rates. Another thing to note is that the coefficients of inflation and growth rates on long-term interest rates are both negative, and so have the opposite signs to their expected values. This is more notable for the latter one, which is larger in its magnitude and statistically significant at the 0.1 level.

Next, we follow Dell'Erba and Sola (2016) in splitting the sample into two subsamples, which are divided by the financial crisis of 2007–08 so that the first semi-annual period that belongs to post-financial crisis era is the December forecast in 2007. Although studying the impact of the crisis on the relationship between fiscal policy and long-term interest rates is not our main goal, it is reasonable and effortless to check for possible structural breaks in the coefficients due to the crisis. We begin by estimating the first two models (from columns 1 and 2) separately for pre- and post-financial crisis eras. Columns 3, 4, 5 and 6 lay out the results.

Columns 3 and 4 show that when the sample is limited to the pre-financial crisis period, the coefficients match the preconceptions one has for them much better. Now fiscal policy does not lose its effectiveness under the time fixed effects[34], both coefficients of inflation and growth

[34] To be precise, primary balance-to-GDP ratio declines by 5 basis points but notably does not become statistically insignificant.

rates have positive signs and the R-squared measure indicates much better fit for the data. The same cannot be said when the two models are estimated for the second subsample (columns 5 and 6). Instead, neither of the fiscal policy measures is significant, the coefficients of inflation and growth rates are back to negative and the R-squared measure has remarkably declined. Our results from the first subsample are similar to those of Dell'Erba and Sola (2016), which is not surprising given that our samples are much alike.

We do not content ourselves with only visibly different coefficients in the two subsamples but instead want to observe whether the difference is statistically significant. Therefore, again following Dell'Erba and Sola (2016), we interact the coefficients of primary balance-to-GDP and public debt-to-GDP ratios with dummy variables that receive value 1 or 0 depending on whether the observation took place before or after the crisis. We first test for the structural breaks in the two coefficients separately (columns 1, 2, 3 and 4) and finally add both interacted coefficients simultaneously to the model (columns 5 and 6). All models are estimated with the full sample. Table 7 lays out the results.

The evidence from table 7 indeed indicates that there are structural breaks in the coefficients of both fiscal policy measures. During the pre-crisis period the long-term interest rates respond to deterioration in primary balances by increasing from four to nine basis points [35], while after the crisis the effect ranges from negative six to positive three basis points, and is not statistically significant. For public debt-to-GDP ratio the pattern is the opposite; during the pre-crisis period the level of gross debt to GDP is irrelevant but after the crisis every percentage point in debt increases long-term interest rates by two basis points. Our results suggest that only the primary balance-to-GDP ratio was relevant for long-term interest rates during the pre-crisis period while the same is true for the public debt-to-GDP ratio during the post-crisis period. Results obtained when the interacted coefficients are estimated separately for both fiscal policy measures remain robust once they are added simultaneously to the model. Our results are quite different from those of Dell'Erba and Sola (2016), who find a structural break only for a deficit-gap variable. Their model specification (factor augmented panel (FAP)) is, however, very different from ours, which means their results cannot be directly compared to ours.

[35] With primary balances, we again confront the problem that the results are not robust under time fixed effects, despite the magnitude of the coefficient is relatively close to that one of estimating the baseline regression with the first subsample (9 and 5 vs. 13 and 8 basis points under time trend parameters and time fixed effects, respectively.

Table 7. Long-term interest rates and fiscal policy – structural breaks with crisis.

Variable	(1)	(2)	(3)	(4)	(5)	(6)
$\boldsymbol{i^S}$	0.523 (0.052)***	0.499 (0.062)***	0.512 (0.055)***	0.503 (0.070)***	0.512 (0.055)***	0.509 (0.073)***
$\boldsymbol{\pi}$	-0.108 (0.097)	-0.109 (0.104)	-0.126 (0.093)	-0.135 (0.102)	-0.121 (0.090)	-0.126 (0.097)
$\boldsymbol{g}$	-0.358 (0.191)#	-0.412 (0.225)#	-0.294 (0.164)#	-0.402 (0.208)#	-0.297 (0.163)#	-0.401 (0.202)#
$\boldsymbol{PB}$			-0.049 (0.029)	-0.006 (0.030)		
$\boldsymbol{PD}$	0.016 (0.008)#	0.012 (0.007)#				
$\boldsymbol{PreC * PB}$	-0.093 (0.042)*	-0.043 (0.046)			-0.089 (0.036)*	-0.049 (0.037)
$\boldsymbol{PostC * PB}$	-0.033 (0.059)	0.035 (0.070)			0.011 (0.066)	0.059 (0.074)
$\boldsymbol{PreC * PD}$			0.004 (0.005)	-0.001 (0.005)	0.005 (0.005)	-0.001 (0.005)
$\boldsymbol{PostC * PD}$			0.020 (0.010)*	0.016 (0.009)#	0.021 (0.010)*	0.017 (0.009)#
Country FE	Yes	Yes	Yes	Yes	Yes	Yes
Time FE	No	Yes	No	Yes	No	Yes
Time trend	Yes	No	Yes	No	Yes	No
$\mathbf{R^2}$	0.742	0.794	0.754	0.802	0.756	0.804
N. of periods	12–55	12–55	12–55	12–55	12–55	12–55
N. of countries	29	29	29	29	29	29
N. of obs.	1281	1281	1281	1281	1281	1281

Notes: (1) Significance codes are the following: 0.1 '#', 0.05 '*', 0.01 '**' & 0.001 '***'. (2) Standard errors (in parenthesis) are corrected for heteroscedasticity by using vce(robust) option in Stata 14.2. (3) 'Country FE', 'Time FE' and 'Time trend' denote the usage of country fixed effects, time fixed effects and linear and quadratic time trends, respectively. (4) 'PreC' and 'PostC' denote the pre- and post-financial crisis dummy variables, respectively.

Source: Author's calculations.

Two conclusions can be drawn based on the results laid out in tables 6 and 7. First, the baseline regression describes the variation of long-term interest rates much better in the first subsample than either in the second one or in the full sample. This indicates that there are omitted factors in the baseline regression that cannot be controlled for with the country and time fixed effects alone, and that these factors play a much more important role in the latter half of the sample. The change in the coefficient signs of both inflation and growth rates implies that the insufficiency of the model may be linked to the low inflation and growth environment that has prevailed since the crisis. The notable drops in both of their levels around the financial crisis (figures 5B and 5C, respectively) combined with the pickups in many governments' long-term interest rates could explain these negative coefficients. One also notices a similar but smaller decline in the coefficient on short-term interest rate. Even though the coefficients on the control variables are not important per se, these finds suggest that the common set of control variables may not be sufficient when the sample also covers the post-financial crisis era.

Second, only one of the two fiscal policy measures seem to have mattered for long-term interest rates at a time; primary balance-to-GDP and public debt-to-GDP ratio during the pre- and post-financial crisis eras, respectively. This finding is plausible once one recalls the severe mispricing of sovereign risk that took place in financial markets during the pre-crisis era since the beginning of 2000s (Beirne & Frantzscher, 2013, 60 & 71). This may have caused the long-term interest rates to be only weakly responsive to the public debt-to-GDP ratio, leading investors to price only the primary balances as an indicator of the course of public sector finances. Once the underlying problems became evident, the pricing shifted towards the evaluation of sovereign risk, which had been overlooked previously.

5.2 Check for nonlinearities

While the above results hint of the reasons behind the inadequacy of the baseline regression with the full sample, they do not give any hard evidence on the underlying factors that caused this insufficiency. For this reason, we continue working with the full sample, without interacting the fiscal policy measures with pre- and post-financial crisis dummy variables, aiming at improving the baseline regression based on the information acquired in section 2.

We first consider whether nonlinearities play an important role in our sample. As was discussed earlier, the nonlinear relationship between fiscal policy and long-term interest rates is justified given default and liquidity premiums embodied in the government bond yields. Our approach to testing for nonlinearities is almost identical to that of Ardagna et al. (2007)[36]: First we add the squared terms of both fiscal policy measures to the baseline regression and then test whether the nonlinearities only played a role when the primary balance-to-GDP and public debt-to-GDP ratios were above their time-varying median values ((MPB) and (MPD), respectively), which

[36] Besides the different samples, the only differences are that Ardagna et al. (2007) employ only time fixed effects and use whole sample means instead of time-varying sample means.

could be the case if investors were strongly responsive to relatively extreme fiscal deterioration. For the latter, the squared deviations between the individual and the time-varying median values are interacted with dummy variables that equal 1 if the fiscal policy measures are above their time-varying median values, and 0 otherwise. Table 8 lays out the results.

Table 8. Long-term interest rates and fiscal policy – nonlinearities.

Variable	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
i^S	0.504 (0.044)***	0.475 (0.057)***	0.531 (0.041)***	0.489 (0.054)***	0.554 (0.056)***	0.518 (0.067)***
π	-0.116 (0.093)	-0.120 (0.092)	-0.126 (0.103)	-0.123 (0.103)	-0.105 (0.105)	-0.117 (0.112)
g	-0.363 (0.189)#	-0.422 (0.215)#	-0.353 (0.188)#	-0.423 (0.230)#	-0.345 (0.180)#	-0.411 (0.218)#
PB	-0.067 (0.037)#	-0.001 (0.041)	-0.108 (0.036)**	-0.040 (0.038)	-0.100 (0.031)**	-0.048 (0.033)
PD	-0.004 (0.014)	-0.017 (0.018)	0.016 (0.004)***	0.006 (0.006)	0.024 (0.010)*	0.017 (0.010)#
PB^2	-0.005 (0.007)	-0.007 (0.006)				
PD^2	0.000 (0.000)	0.000 (0.000)				
$(PB - MPB)^2 * DumPB$			-0.022 (0.006)***	-0.011 (0.008)		
$(PD - MPD)^2 * DumPD$			-0.000 (0.000)	0.000 (0.000)		
$(PB - MPB)^2 * DumPD$					-0.023 (0.011)*	-0.019 (0.013)
$(PD - MPD)^2 * DumPB$					-0.000 (0.000)*	-0.000 (0.000)
Country FE	Yes	Yes	Yes	Yes	Yes	Yes
Time FE	No	Yes	No	Yes	No	Yes
Time trend	Yes	No	Yes	No	Yes	No
$\mathbf{R^2}$	0.746	0.800	0.745	0.794	0.749	0.796
N. of periods	12–55	12–55	12–55	12–55	12–55	12–55
N. of countries	29	29	29	29	29	29
N. of obs.	1281	1281	1281	1281	1281	1281

Notes: (1) Significance codes are the following: 0.1 ‘#’, 0.05 ‘*’, 0.01 ‘**’ & 0.001 ‘***’. (2) Standard errors (in parenthesis) are corrected for heteroscedasticity by using vce(robust) option in Stata 14.2. (3) ‘Country FE’, ‘Time FE’ and ‘Time trend’ denote the usage of country fixed effects, time fixed effects and linear and quadratic time trends, respectively. (4) ‘DumPB’ and ‘DumPD’ denote dummy variables that equal 1 if the primary balance-to-GDP and public debt-to-GDP ratios are above their time-varying median values, and 0 otherwise. (5) MPB and MPD are the time-varying median primary balance-to-GDP and public debt-to-GDP ratios, respectively.

Source: Author’s calculations.

The results from columns 1 and 2 show no evidence for the nonlinear relationship between fiscal policy and long-term interest rates as the magnitudes of the squared terms of the coefficients on primary balance-to-GDP and public debt-to-GDP ratios are very small and not statistically significant. For primary balances, the coefficient even has the opposite sign to its expected value. Adding the interacted squared deviations between the individual and time-varying median values (columns 3, 4, 5 and 6) yield the same conclusion – the only significant results, from columns 3 and 5, have opposite signs to their expected values and their significance quickly disappears under the time fixed effects, meaning that the results are not robust to controlling for the time-constant global factors. Hence, we conclude that there is no evidence for nonlinearities in our sample.

Our results contradict those of Baldacci and Kumar (2010) and Ardagna et al. (2007) who both find clear evidence on nonlinearities. The key difference between the results may again reside in the time spans of the employed samples; the sample used by Baldacci and Kumar (2010) covers periods from 1980 to 2007 while Ardagna et al. (2007) utilize one from 1960 to 2002. This means that neither of the referenced studies covers the post-financial crisis era, which makes a large part of our sample. Indeed, our results are much closer to those of Dell'Erba and Sola (2016), who find only weak evidence on nonlinearities by employing a sample from 1989 to 2013.

5.3 Check for the influence of the omitted factors

Next, we extend the model to tackle the influence from the presence of the UMPs and also consider the possible effects from the default and liquidity premiums more directly than by adding the nonlinear variants of the fiscal policy measures to the model. The addition of these variables leaves financial openness as the last factor that is omitted in the baseline regression, but to whom addition there is a strong reasoning based on our theoretical framework. However, for now we shall skip over the topic of financial openness as it is considered in more detail in the following subsection.

As discussed earlier, short-term interest rates may be insufficient at depicting the true stance of monetary policy when they are at or close to their effective lower bound and other methods, such as UMPs, have been put into operation to further depress the level of interest rates. During the past decade, many central banks, such as Bank of England, Bank of Japan, European Central Bank, Sveriges Riksbank, Swiss National Bank and the Federal Reserve System, have been faced with this type of situation as they have been trying to combat the economic downturn since the financial crisis of 2007–08. Consequently, our data includes many observations where UMPs were an important part in the monetary policy stances of the representative central banks. Ignoring their influence on long-term interest rates could hence bias our results. This could also explain the remarkable decline in the coefficient of the short-term interest rates when the sample is limited to the second subsample.

We address the possible effect from the presence of UMPs indirectly by noting the country-period combinations with central bank policy rates set at 0.5 % or below[37], which arguably is a level with very little room to further reduce the policy rate (Eggertsson & Woodford, 2003, 139). In practice, this is done by constructing a dummy variable ($LowCBR$) based on the central bank policy rates (CBR) that equals 1 if the prevailing policy rate is 0.5 % or lower, and 0 otherwise. In other words, ($LowCBR$) is meant to proxy the periods when UMPs were in operation to further decrease long-term interest rates. An indirect approach was chosen as we do not want to distinguish between the effects of different types of UMPs but only want to consider for the effect of their presence on long-term interest rates.

We also add sovereign credit ratings (CR) that prevailed during each period of time in order to better capture the default premium embodied in the long-term interest rates. If financial markets priced the public indebtedness correctly, the nonlinear variants of the fiscal policy measures should capture the impact of the default premium on long-term interest rates, at least to some extent, which would make the addition of sovereign credit ratings to the model less interesting. However, as was seen in the last subsection, this is not the case in our sample. This is not surprising once one recalls the severe mispricing of sovereign risk that took place during the pre-crisis period 2000–07 (Beirne & Fratzscher, 2013, 60 & 71). Figure 11 illustrates a possible explanation why the nonlinear variants of the fiscal policy measures may fail to accurately describe the impact of the default premium on long-term interest rates in our sample by plotting both credit rating changes and median values of the public debt-to-GDP ratios in one chart. As one can see, there are clear differences between the developments of the two variables;

[37] Out of the 1281 observations, 216 are such were the respective central banks had set their policy rates at or below 0.5 %. Appendix C plots these observations for each country against time.

sovereign credit ratings were upgraded more rapidly than the public debt-to-GDP ratios declined during the early 2000s, but also their downgrading was more intense than the incurring of public debt after the crisis.

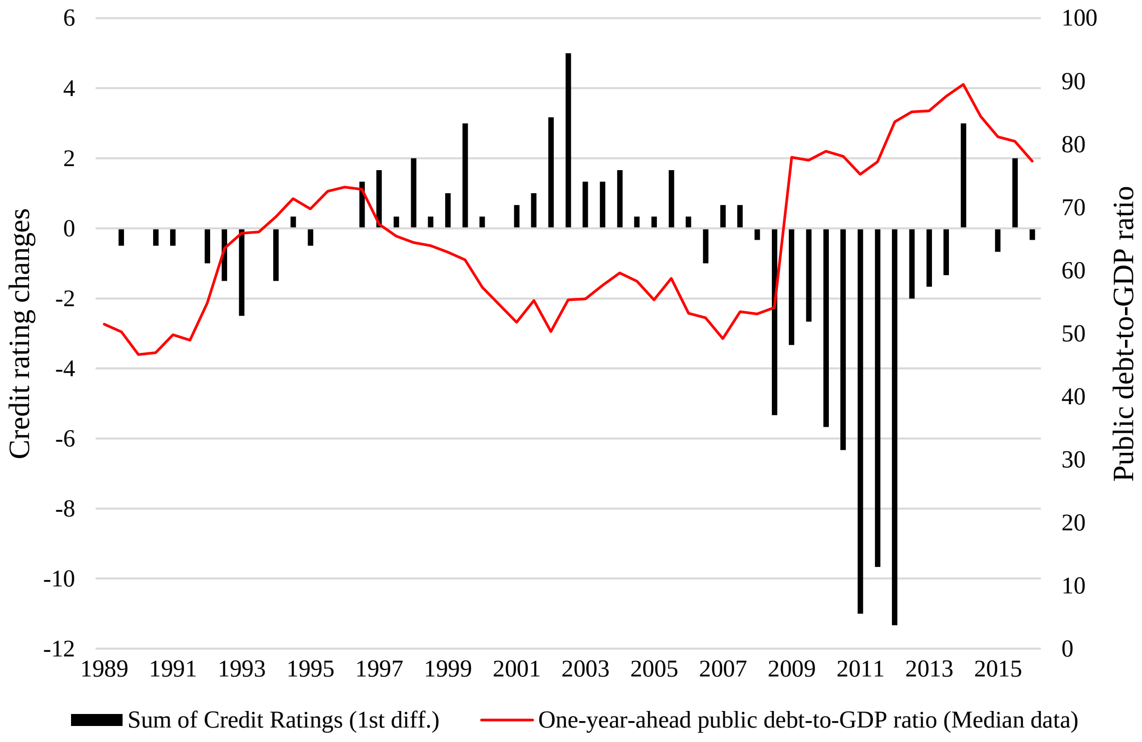


Figure 11. Credit rating changes and public debt-to-GDP ratios.

Finally, we also construct two dummy variables, $(SmallGBM)$ and $(RSmallGBM)$, based on the government bond market sizes (GBM), to account for possible raises in government bond yields due to illiquid government bond markets. The prior equals 1 for observations with public debt market capitalization less than 200 billion dollars, and 0 otherwise, whereas the latter equals 1 for observations which for the size of the government bond market is below the time-varying sample median, and 0 otherwise. Again, the nonlinear variants of the fiscal policy measures could capture this effect (Ardagna et al., 2007, 8), but for samples which consist of very different sized economies, such as ours, there is an obvious flaw with this type of specification: small and large economies reach the threshold of a liquid government bond market (which was estimated to lie around from 100 to 200 billion dollars by McCauley and Remolona (2000)) with very different public debt-to-GDP ratios.

We begin by adding both (CR) and $(LowCBR)$ simultaneously to the baseline regression (columns 1 and 2) and then add the two proxy variables of the illiquid government bond market,

($SmallGBM$) and ($RSmallGBM$), to the model, one at a time (columns 3, 4, 5 and 6). Table 9 lays out the results.

Table 9. Long-term interest rates and fiscal policy – omitted factors.

Variable	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
$\boldsymbol{i^S}$	0.491 (0.034)***	0.479 (0.040)***	0.486 (0.033)***	0.470 (0.039)***	0.488 (0.036)***	0.476 (0.044)***
$\boldsymbol{\pi}$	-0.020 (0.102)	-0.023 (0.105)	0.002 (0.114)	-0.001 (0.114)	-0.016 (0.106)	-0.019 (0.108)
$\boldsymbol{g}$	-0.213 (0.110)#	-0.247 (0.128)#	-0.212 (0.108)#	-0.244 (0.123)#	-0.218 (0.110)#	-0.254 (0.128)#
$\boldsymbol{PB}$	-0.132 (0.033)***	-0.067 (0.027)*	-0.128 (0.034)***	-0.060 (0.028)*	-0.128 (0.033)***	-0.062 (0.027)*
$\boldsymbol{PD}$	-0.004 (0.006)	-0.009 (0.006)	-0.004 (0.005)	-0.008 (0.006)	-0.004 (0.006)	-0.009 (0.006)
$\boldsymbol{CR}$	-0.512 (0.200)*	-0.511 (0.190)*	-0.505 (0.182)**	-0.507 (0.171)**	-0.510 (0.199)*	-0.512 (0.188)*
$\boldsymbol{LowCBR}$	-1.133 (0.333)**	-0.670 (0.248)*	-1.085 (0.304)***	-0.596 (0.211)**	-1.100 (0.315)**	-0.634 (0.232)*
$\boldsymbol{SmallGBM}$			-0.848 (0.737)	-0.872 (0.784)		
$\boldsymbol{RSmallGBM}$					-0.465 (0.304)	-0.433 (0.337)
Country FE	Yes	Yes	Yes	Yes	Yes	Yes
Time FE	No	Yes	No	Yes	No	Yes
Time trend	Yes	No	Yes	No	Yes	No
R^2	0.791	0.835	0.789	0.835	0.787	0.833
N. of periods	12–55	12–55	12–55	12–55	12–55	12–55
N. of countries	29	29	29	29	29	29
N. of obs.	1281	1281	1266	1266	1266	1266

Notes: (1) Significance codes are the following: 0.1 '#', 0.05 '*', 0.01 '**' & 0.001 '***'. (2) Standard errors (in parenthesis) are corrected for heteroscedasticity by using vce(robust) option in Stata 14.2. (3) 'Country FE', 'Time FE' and 'Time trend' denote the usage of country fixed effects, time fixed effects and linear and quadratic time trends, respectively. (4) 'LowCBR', 'SmallGBM', and 'RSmallGBM' denote dummy variables that equal one if the central bank rate is set at 0.5 % or below, the size of the government bond market is below $200 billion, and the size of the government bond market is below the time-varying sample median, respectively.

Source: Author's calculations.

The results suggest that the presence of UMPs are an important determinant of the long-term interest rates in our sample; its impact ranges from -67 to -113 basis points and is highly significant under both specifications. On additional note, the effect does not overlap with that of short-term interest rates. Instead, the coefficient of short-term interest rates remains very close to the 50 basis points area, which is around the level estimated in the baseline regression with the full sample (see table 6, columns 1 and 2). Credit ratings prove to be an important factor for the long-term interest rates as well; the impact from a one-notch upgrade in the average credit rating is a solid 51 basis points under both specifications.

While both of these variables prove to be important determinants of the long-term interest rates in our model, our actual interest lies in how the two fiscal policy measures perform under this specification. Beginning with the flow variable, a one percentage point deterioration in the primary balance-to-GDP ratio increases the long-term interest rates from 7 to 13 basis points, meaning that under a more robust set of control variables the effect is 6 basis points higher than in the baseline regression. Most importantly, the result now proves to be robust switching from time trend parameters to the time fixed effects, despite the 6 basis points decline in the coefficient. Interestingly enough, the magnitude of the effect actually corresponds near to one-to-one with that of the baseline regression with the first subsample. This implies that the addition of (CR) and ($LowCBR$) to the model reveals the part of the impact of the primary balance-to-GDP ratio on long-term interest rates, which is masked in the baseline regression with both the full sample and the second subsample.

For the stock measure the effect is close to zero and not statistically significant. The fact that the one basis point effect (table 6, columns 1, 2, 3 and 4) disappears completely when the credit ratings are added to the model implies that bulk of the effect of the public debt-to-GDP ratio comes over to long-term interest rates via the default premium, and once this part of the effect is controlled for with another measure, the public debt-to-GDP ratio becomes irrelevant for long-term interest rates.

The remainder of the effect of the public debt-to-GDP ratio has a negative coefficient, which could be due to a liquidity effect (Ardagna et al., 2007, 8). However, the results from adding either of the two proxy variables of an illiquid government bond market (columns 3, 4, 5 and 6) suggest that the liquidity premium does not have a statistically significant impact on the long-term interest rates, or alternatively, that our specification fails at revealing it in our sample. Either way, we conclude that there is no evidence that the fiscal policy would have had an effect on long-term interest rates via the liquidity premium.

While the results acquired in this subsection are, to a great extent, plausible, they are unfortunately not comparable with those of other studies as we are, to the knowledge of the author, the first ones to employ such measures in this literature. The reason why other researchers have not considered the addition of these measures may again be linked to the fact that their studies lack the coverage of the post-financial crisis period, which arguably is the era that complicates the results of the baseline regression. Simply put, the addition of these measures may not have been needed when the impact of fiscal policy on long-term interest rates has been examined with pre-crisis data alone.

5.4 Check for the influence of financial openness

Both Ricardian equivalence proposition and high degree of financial openness could justify the lack of a response of interest rates to fiscal deterioration (Clayes et al., 2012, 57). Despite the clear evidence on the effect, we are interested in testing whether the magnitude of the effect differs for economies with different degrees of financial openness as is suggested by our theoretical framework.

We follow Aisen and Hauner (2013) in their methodology in analysing the influence of financial openness on the impact of fiscal policy on long-term interest rates. First, we construct two sets of dummy variables, $(LowCO)$ and $(HighCO)$, and $(LowCO2)$ and $(HighCO2)$, so that the 'High' and 'Low' dummies equal one for values above or equal and zero for values below the time-varying sample median values, respectively. The two sets of dummy variables differ from each other only by the data that were used for the construction of the dummy variables: 2017 July update of the index of capital account openness by Chinn and Ito (2006), $(KAOPEN)$, was used for the prior pair and the self-calculated sums of foreign assets and liabilities over GDP, $(FAFLGDP)$, for the latter. Next, we interact the constructed sets with both fiscal policy measures so that we receive a total of two groups of four interacted fiscal policy variables. Both groups are then added to the model that includes (CR) and $(LowCBR)$[38], one at a time. The reason for testing the influence of financial openness with two alternative

[38] We continue working with the model that includes (CR) and $(LowCBR)$ as control variables (table 9, columns 1 and 2), as they proved to be important determinants of long-term interest rates in our sample.

measurements is simply that it allows us to check for the robustness of the results. Table 10 reports the results.

Table 10. Long-term interest rates and fiscal policy – financial openness.

Variable	(1)	(2)	(3)	(4)
i^S	0.479 (0.035)***	0.602 (0.044)***	0.487 (0.032)***	0.470 (0.038)***
π	0.001 (0.110)	-0.041 (0.116)	-0.018 (0.101)	-0.019 (0.098)
g	-0.184 (0.106)#	-0.180 (0.127)	-0.219 (0.114)#	-0.263 (0.129)#
$LowCO * PB$	-0.158 (0.063)*	-0.116 (0.060)#		
$HighCO * PB$	-0.122 (0.031)***	-0.064 (0.026)*		
$LowCO * PD$	-0.005 (0.008)	-0.011 (0.008)		
$HighCO * PD$	-0.005 (0.005)	-0.010 (0.005)*		
$LowCO2 * PB$			-0.146 (0.048)**	-0.084 (0.037)*
$HighCO2 * PB$			-0.113 (0.030)***	-0.029 (0.036)
$LowCO2 * PD$			-0.004 (0.005)***	-0.007 (0.004)#
$HighCO2 * PD$			-0.006 (0.008)	-0.012 (0.009)
CR	-0.546 (0.211)*	-0.552 (0.192)**	-0.512 (0.198)*	-0.512 (0.186)**
$LowCBR$	-1.095 (0.350)**	-1.002 (0.312)**	-1.131 (0.331)**	-0.650 (0.237)**
Country FE	Yes	Yes	Yes	Yes
Time FE	No	Yes	No	Yes
Time trend	Yes	No	Yes	No
R^2	0.781	0.812	0.791	0.836
N. of periods	10–53	10–53	12–55	12–55
N. of countries	28	28	29	29
N. of obs.	1195	1195	1281	1281

Notes: (1) Significance codes are the following: 0.1 '#', 0.05 '*', 0.01 '**' & 0.001 '***'. (2) Standard errors (in parenthesis) are corrected for heteroscedasticity by using vce(robust) option in Stata 14.2. (3) 'Country FE', 'Time FE' and 'Time trend' denote country fixed effects, time fixed effects and time trend parameters, respectively. (4) 'LowCBR', 'LowCO' ('LowCO2'), and 'HighCO' ('HighCO2') denote dummies that equal one if central bank rate is set at 0.5 % or below, financial openness is below, and above or equal to the time-varying sample medians, respectively.

Source: Author's calculations.

Beginning with the prior specification (columns 1 and 2), the results indicate that the effect from a one percentage point deterioration in primary balance-to-GDP ratio on long-term interest rates is from four to six basis points lower for economies with high capital account openness. The results are statistically significant and robust for switching from time trend parameters to time fixed effects. The same cannot be said for the public debt-to-GDP ratio though, as it proves to be insignificant for both economies with low and high capital account openness.

The latter specification provides very similar results to the first one (columns 3 and 4). Although the coefficients of primary balance-to-GDP ratios are now a couple of basis points lower for each variant of the measurement, the difference in the magnitude of the effect for economies with low and high capital account openness is almost identical, from four to five basis points, to that of the first specification. The only notable difference here is that the effect from deterioration in primary balances becomes insignificant for economies with high capital account openness under the time fixed effects. The difference between the two specifications for the other fiscal policy measurement is that public debt-to-GDP ratio is now found to be significant for economies with low capital account openness, although the coefficients are again almost identical to those of the first specification. Despite their statistical significance, their economic significance is still far away from that of primary balance-to-GDP ratio. The negative coefficients of the different variants of public debt-to-GDP ratio again imply that the liquidity effect could have a minor impact on long-term interest rates, even though this could not be revealed in the prior subsection with our specification of choice.

We conclude that the above evidence suggests that financial openness, measured as capital account openness, influences how strongly fiscal deterioration comes over to long-term interest rates. Our results are similar to those of Aisen and Hauner (2013) with one remarkable difference: they find the effect from budget deficits on interest rates to be 67 and two basis points for economies with low and high capital account openness, respectively, by employing the very same index of capital account openness by Chinn and Ito (2006) that is used for the first specification in this subsection. Hence, their estimated effect of low capital account openness on the impact of fiscal deterioration on interest rates is 65 basis points, while our results suggests that the effect is only from four to six basis points. The difference between the results of Aisen and Hauner (2013) and ours may be due to the countries included in the samples: Aisen and Hauner (2013) employ a much wider dataset with a total of 60 advanced and emerging market economies (table 1), whereas we include only 29 OECD economies in our sample. Consequently, their sample includes economies with much more divergent degrees of capital account openness compared to our sample that only includes OECD members, from which most have relatively high degrees of capital account openness. Hence, instead of being a major complication for our results, the fact that Aisen and Hauner (2013) find the influence of capital account openness to the impact of fiscal policy on interest rates to be much higher is very much reasonable.

6 Conclusions

In this thesis we studied the effects of fiscal policy on long-term interest rates by applying fixed effects estimation on a panel of 29 OECD economies over the last three decades. We identified the different channels through which fiscal policy can be expected to influence interest rates and reviewed the empirical literature as to build up the prerequisite information for our empirical analysis. The actual empirical assessment started off with an estimation of a so-called baseline regression with common specifications that provided the base case for our analysis. The model was then expanded in several ways to tackle factors that were omitted in the baseline regression but had strong theoretical justification to be included in the model. Consequently, we found the presence of UMPs, credit ratings and capital account openness to have an important effect on the relationship between fiscal policy and long-term interest rates.

The theoretical framework of the subject is manifold. Not only does the interest rate effect of fiscal policy occur via various different channels, i.e. national savings as well as inflation, default and liquidity premiums with partly contrary effects, but the effect is also further affected by openness of an economy and stance of monetary policy. This leads the overall interest rate effect of fiscal policy to be dependent on host of elasticities of the different channels, which largely complicates its assessment.

The baseline regression was found to describe poorly the variation of long-term interest rates with our full sample which also covers the post-financial crisis era, but to provide plausible results that are like those obtained in most of the literature with the pre-crisis sample. With the full sample, one percentage point deterioration in primary balance-to-GDP and public debt-to-GDP ratios increased long-term interest rates by seven and two basis points, respectively, whereas with the pre-crisis sample the equivalent effects were 13 and one basis points. The disconcerting find in the results with our full sample were not that the effects matched worse with those of the prior literature, but that they were not robust to switching between our two alternative specifications and produced remarkably poorer fit for the data than those estimated for the pre-crisis sample. Altogether, our evidence suggested that the baseline regression is inadequate with our full sample and that its inadequacy is linked to certain complexities due to the crisis, e.g. the harsh economic development and the changed pricing of sovereign risk, which cannot be tackled by the baseline regression alone.

Against this background, we sought for a way to improve the fit of our model for the full sample based on the information of our theoretical framework. We found no evidence that the effect of fiscal policy would be nonlinear, which could, and has previously been found to, improve the model by better capturing default and liquidity premiums embodied in government bond yields than a linear model. A plausible, albeit hypothetical, explanation for the nonexistence of nonlinearities in our sample draws again from its timeline: the severe mispricing of sovereign risk that took place before the crisis (Beirne & Fratzscher, 2013, 60 & 71) and the increased price sensitivity to it since the wake of the crisis create a timespan with two different types of responses to changes in sovereign risk due to fiscal expansion, which may cause nonlinearities to fail to tackle the interest rate effect of fiscal policy through the default premium.

We extended the literature by considering whether the addition of the presence of UMPs, credit ratings, and government bond market sizes affect the relationship between fiscal policy and long-term interest rates, and so would correct the shortcomings of the baseline regression with our full sample. We found the prior two to be important determinants of long-term interest rates in our full sample. Moreover, their addition to the model altered the interest rate effects through both fiscal policy measures so that they better matched with those of the baseline regression with the pre-crisis sample. Our results implied that they successfully tackled the majority of the complexities due to the crisis and so provided a more robust set of control variables for studying the interest rate effect of fiscal policy with post-crisis data. Additionally, we found credit ratings to crowd out the interest rate effect of public debt-to-GDP ratio, implying that the bulk of its effects come over to long-term interest rates via the default premium, and that when the default premium is controlled with an alternative measurement, the stock of debt hardly plays a role on long-term interest rates anymore. Despite its negative residual effect rooted for an impact through liquidity premium, we found no evidence for it with the help of government bond market sizes.

Finally, we found clear evidence that for more financially open economies the interest rate effect of fiscal policy is smaller than for less financially open economies. The magnitude of our finding was, however, remarkably lower than in the prior literature. The fact that our sample includes only OECD economies that all have relatively high degrees of financial openness means that also the differences between the degrees are smaller than in the sample of Aisen and Hauner (2013) who employ a much wider dataset with both advanced and emerging market economies. Hence, the results with a smaller magnitude are not problematic, but in fact make sense.

The results of the thesis are plausible from a theoretical point of view and line up very well with those of the earlier literature to the extent that our work is comparable with theirs. The fiscal policy comes over to long-term interest rates mainly through the flow variable, i.e. the primary balance-to-GDP ratio, whereas the effect through the stock variable, i.e. the public debt-to-GDP ratio, is much weaker, if existent. The magnitude of both effects also match closely to those of the prior literature. The obvious drawback of our results is that our proposed framework addressing the complexities due to the crisis is not comparable to other studies, as we are the first to employ the presence of UMPs and credit ratings in this literature. The reason for this may again reside in our sample. The empirical literature on the topic lacks the coverage of the post-financial crisis era. To our knowledge, only one paper exists with a sample timeline that reaches our current decade. As the baseline regression, which is commonly used in the literature, provides plausible results with the pre-crisis data, the addition of these measures may simply not have been needed before.

A natural next step for further research would be to consolidate our findings with more elaborate econometric methodologies. Employing a panel vector autoregressive model, for instance, would allow to address interdependencies between independent variables and to evaluate the cumulative interest rate effect of fiscal policy over time, and so could both reassert the results of the thesis as well as provide brand new information on the topic. While the thesis provided evidence that the public debt-to-GDP ratio mainly influences interest rates through the default premium, the effect through the remainder of the channels remained vague. It would be interesting to study the different channels more in detail, as a clarification on their intensity would greatly benefit this literature. For example, one could cast light on the strength of the interest rate effect via national savings by combining World Bank data on savings rates with the type of dataset employed in the thesis. Finally, limiting the sample of countries to those with a common currency, *i.e. to Eurozone countries*, could be reasonable, as this would not only allow to exclude the effect of exchange rates on long-term interest rates, but also that of changes in the exchange-rate regimes. This would also ease the control of cyclical conditions, as the Eurozone countries share common monetary policy and inflation dynamics. The drawback of this approach is, however, the significant decline in the sample size, which could hinder the generalization of the results.

References

Aisen, A. & Hauner, D. (2013). Budget deficits and interest rates: a fresh perspective. *Applied Economics*, Vol. 45, Iss. 17, 2501–2510.

Aizenman, J., Hutchison, M. & Jinjarak, Y. (2013). What is the risk of European sovereign debt defaults? Fiscal space, CDS spreads and market pricing of risk. *Journal of International Money and Finance*, Vol. 34, 37–59.

Ardagna, S. (2009). Financial markets' behaviour around episodes of large changes in the fiscal stance. *European Economic Review*, Vol. 53, Iss. 1, 37–55.

Ardagna, S., Caselli, F. & Lane, T. (2007). Fiscal Discipline and the Cost of Public Debt Service: Some Estimates for OECD Countries. *The B.E. Journal of Macroeconomics*, Vol. 7, Iss. 1, Art. 28, 1–33.

Arezki, R., Candelon, B. & Sy, A. (2011). Sovereign Rating News and Financial Markets Spillovers: Evidence from the European Debt Crisis. *CESifo Working Paper*, No. 3411, 1–28.

Artis, M. & Marcellino, M. (2001). Fiscal forecasting: The track record of the IMF, OECD and EC. *Econometrics Journal*, Vol. 4, Iss. 1, 20–36.

Auerbach, A. J. & Kotlikoff, L. J. (1987). *Dynamic Fiscal Policy* (1^{st} ed.). Cambridge: Cambridge University Press.

Baldacci, E. & Kumar, M. S. (2010). Fiscal Deficits, Public Debt, and Sovereign Bond Yields. *IMF Working Paper*, No. 10/184, 1–28.

Bank of England. (1999). *The transmission mechanism of monetary policy*. http://www.bankofengland.co.uk/publications/Documents/other/ monetary/ montrans.pdf. July 2017.

Barro, R. J. (1974). Are Government Bonds Net Wealth? *Journal of Political Economy*, Vol. 82, No. 6, 1095–1117.

Beck, T., Demirgüc-Kunt, A. & Levine, R. (2000). A New Database on the Structure and Development of the Financial Sector. *The World Bank Economic Review*, Vol. 14, Iss. 3, 597–605.

Beetsma, R. & Massimo, G. (2010). Discretionary fiscal policy: Review and estimates for the EU. *CESifo Working Paper*, No. 2948, 1–28.

Beirne, J. & Fratzscher, M. (2013). The pricing of sovereign risk and contagion during the European sovereign debt crisis. *Journal of International Money and Finance*, Vol. 34, 60–82.

Bernheim, B. D. (1989). A Neoclassical Perspective on Budget Deficits. *Journal of Economic Perspectives*, Vol. 3, Number 2, 55–72.

Bervas, A. (2006). Market liquidity and its incorporation into risk management. *Financial Stability Review Banque de France*, No. 8, 63–79.

Boothe, P. M. & Reid, B. G. (1989). Asset returns and government budgets in a small open economy: Empirical evidence for Canada. *Journal of Monetary Economics*, Vol. 23, Iss. 1, 65–77.

Cantor, R. & Packer, F. (1996). Determinants and Impacts of Sovereign Credit Ratings. *FRBNY Economic Policy Review*, Vol. 2, Iss. 2, 37–53.

Canzoneri, M. B., Cumby, R. E. & Diba, B. T. (2002). Should the European Central Bank and the Federal Reserve Be Concerned about Fiscal Policy? *Working Paper*, presented at Federal Reserve Bank of Kansas City's symposium on Rethinking Stabilization Policy, Jackson Hole, 333–389.

Catão, L. A. V. & Terrones, M. E. (2005). Fiscal deficits and inflation. *Journal of Monetary Economics*, Vol. 52, Iss. 3, 529–554.

Chinn, M. D. & Frankel, J. (2007). Debt and Interest Rates: The U.S. and the Euro Area. *Economics Discussion Papers*, No. 2007–11, 1–26.

Chinn, M. D. & Ito, H. (2006). What matters for financial development? Capital controls, institutions, and interactions. Journal of Development Economics, Vol. 81, Iss. 1, 163–192.

Cimadomo, J. (2012). Fiscal Policy in Real Time. *Scandinavian Journal of Economics*, Vol. 114, Iss. 2, 409–439.

Clayes, P., Moreno, R. & Suriñach, J. (2012). Debt, interest rates, and integration of financial markets. *Economic Modelling*, Vol. 29, Iss. 1, 48–59.

Del Negro, M., Giannoni, M. & Patterson, C. (2012). The forward guidance puzzle. *Staff Report*, Federal Reserve Bank of New York, No. 574, 1–24.

Dell'Erba, S. & Sola, S. (2016). Does fiscal policy affect interest rates? Evidence from a factor-augmented panel. *The B.E. Journal of Macroeconomics*, Vol. 16, Iss. 2, 395–437.

Dellas, H., Neusser, K. & Wälti, M. (2005). Fiscal Policy in Open Economies. *Working Paper*, Department of Economics, University of Bern, 1–35.

Diamond, P. A. (1965). National Debt in a Neoclassical Growth Model. *The American Economic Review*, Vol. 55, No. 5, 1126–1150.

Eggertsson, G. B. & Woodford, M. (2003). The Zero Bound on Interest Rates and Optimal Monetary Policy. *Brookings Papers on Economic Activity*, Vol. 2003, No. 1, 139–211.

Elmendorf, D. W. & Mankiw, N. G. (1999). Government Debt. In *Handbook of Macroeconomics, Volume 1C* (1st ed.). Amsterdam: Elsevier Science B.V.

Engen, E. M. & Hubbard, R. G. (2004). Federal Government Debt and Interest Rates. *NBER Working paper*, No. 10681, 83–137.

European Commission. (2017). *Stability and convergence programmes*. https://ec.europa.eu/info/business-economy-euro/economic-and-fiscal-policy-coordination/eu-economic-governance-monitoring-prevention-correction/stability-and-growth-pact/stability-and-convergence-programmes_en. July 2017.

Evans, P. (1987a). Do budget deficits raise nominal interest rates?: Evidence from six countries. *Journal of Monetary Economics*, Vol. 20, Iss. 2, 281–300.

Evans, P. (1987b). Interest Rates and Expected Future Budget Deficits in the United States. *Journal of Political Economy,* Vol. 95, No. 1, 34–58.

Faini, R. (2006). Fiscal policy and interest rates in Europe. *Economic Policy*, Vol. 21, Iss. 47, 443–489.

Fawley, B. W. & Neely, C. J. (2013). Four Stories of Quantitative Easing. Federal Reserve Bank of St. Lois Review, Vol. 95, No. 1, 51–88.

Friedman, M. (1957). The Permanent Income Hypothesis. In *A Theory of the Consumption Function* (1st ed.). New Jersey: Princeton University Press.

Gale, W. G. & Orszag, P. T. (2004). Budget Deficits, National Saving, and Interest Rates. *Brookings Papers on Economic Activity*, Iss. 2, 101–210.

Gómez-Puig, M. (2006). Size matters for liquidity: Evidence from EMU sovereign yield spreads. *Economics Letters*, Vol. 90, Iss. 2, 156–162.

Gruber, J. W. & Kamin, S. B. (2012). Fiscal Positions and Government Bond Yields in OECD Countries. *Journal of Money, Credit and Banking*, Vol. 44, Iss. 8, 1563–1587.

Haugh, D., Ollivaud, P. & Turner, D. (2009). What Drives Sovereign Risk Premiums?: An Analysis of Recent Evidence from Euro Area. *OECD Economics Department Working Papers*, No. 718, 1–24.

Hsiao, C. (2003). *Analysis of Panel Data* (2nd ed.). New York: Cambridge University Press.

Im, K. S., Pesaran, M. H. & Shin, Y. (2003). Testing for unit roots in heterogenous panels. *Journal of Econometrics*, Vol. 115, Iss. 1, 53–74.

Inoue, H. (1999). The Structure of Government Securities Markets in G10 Countries: Summary of Questionnaire Results. *BIS CGFS Papers*, No. 11, 1–22.

Joyce, M., Miles, D., Scott, A. & Vayanos, D. (2012). Quantitative Easing and Unconventional Monetary Policy – an Introduction. The Economic Journal, Vol. 122, Iss. 564, F271–F288.

Kennedy, P. (2003). *A Guide to Econometrics* (5th ed.). Oxford: Blackwell Publishing.

Krishnamurthy, A. & Vissing-Jorgensen, A. (2011). The Effects of Quantitative Easing on Interest Rates: Channels and Implications for Policy. *Brookings Papers on Economic Activity*, Fall, 215–287.

Lane, P. R. & Milesi-Ferretti, G. M. (2003). International financial integration. *IMF Staff Papers*, Vol. 50, 82–113.

Lane, P. R. & Milesi-Ferretti, G. M. (2007). The external wealth of nations mark II: Revised and extended estimates of foreign assets and liabilities, 1970–2004. *Journal of International Economics*, Vol. 73, Iss. 2, 223–250.

Laubach, T. (2009). New Evidence on the Interest Rate Effects of Budget Deficits and Debt. *Journal of the European Economic Association*, Vol. 7, Iss. 4, 858–885.

Lin, H.-Y. & Chu, H.-P. (2013). Are fiscal deficits inflationary? *Journal of International Money and Finance*, Vol. 32, 214–233.

Manasse, P. & Roubini, N. (2009). "Rules of thumb" for sovereign debt crises. *Journal of International Economics*, Vol. 78, Iss. 2, 192–205.

Mankiw, N. G. (2012). *Macroeconomics* (8th ed.). New York: Worth Publishers.

Mankiw, N. G. & Miron, J. A. (1986). The Changing Behaviour of the Term Structure of Interest Rates. *The Quarterly Journal of Economics*, Vol. 101, No. 2, 211–228.

Mátyás, L. (2008). *The Econometrics of Panel Data* (3rd ed.). Berlin: Springer.

McCauley, R. & Remolona, E. (2000). Size and liquidity of government bond markets. *BIS Quarterly Review*, November, 52–60.

Mishkin, F. S. (2004). *The Economics of Money, Banking and Financial Markets* (7th ed.). USA: Pearson Addison Wesley.

OECD. (2005). Glossary of statistical terms. https://stats.oecd.org/glossary/detail.asp?ID=6789. May 2017.

OECD. (2016). *OECD Economic Outlook No. 100 (Edition 2016/2)*. http://www.oecd-ilibrary.org/economics/data/oecd-economic-outlook-statistics-and-projections/oecd-economic-outlook-no-100-edition-2016-2_7fa317bf-en?isPartOf=/content/datacollection/eo-data-en. July 2017.

Persyn, D. & Westerlund, J. (2008). Error-correction–based cointegration tests for panel data. *The Stata Journal*, Vol. 8, No. 2, 232–241.

Phillips, P. C. B. & Moon, H. R. (1999). Linear Regression Limit Theory for Nonstationary Panel Data. *Econometrica*, Vol. 67, No. 5, 1057–1111.

Pill, H. & Reichlin, L. (2017). Non-Standard Monetary Policy and Financial Stability: Developing an Appropriate Macro-Financial Policy Mix. *Cambridge University Press*, Forthcoming.

Plosser, C. I. (1982). Government financing decisions and asset returns. *Journal of Monetary Economics*, Vol. 9, Iss. 3, 325–352.

Plosser, C. I. (1987). Fiscal policy and the term structure. *Journal of Monetary Economics*, Vol. 20, Iss. 2, 343–367.

Romer, D. (2000). Keynesian Macroeconomics without the LM Curve. *Journal of Economic Perspectives*, Vol. 14, No. 2, 149–169.

Romer, D. (2012). Advanced macroeconomics (4th ed.). New York: McGraw-Hill.

Solow, R. M. (1956). A Contribution to the Theory of Economic Growth. *The Quarterly Journal of Economics,* Vol. 70, No. 1, 65–94.

Sargent, T. J. & Wallace, N. (1981). Some Unpleasant Monetarist Arithmetic. *Federal Reserve Bank of Minneapolis Quarterly Review,* Vol. 5, No. 3, 1–17.

StataCorp. (2015). *Stata: Longitudinal-Data/Panel-Data Reference Manual* (Rel. 14). College Station: Stata Press.

S&P Global Ratings. (2014). *Criteria | Governments | Sovereigns: Sovereign Rating Methodology*. https://www.standardandpoors.com/en_US/web/guest/article/-/view/type/HTML/id/1813072. July 2017.

S&P Global Ratings. (2017). https://www.standardandpoors.com/en_AP/web/guest/home. August 2017.

Thomas, L. B. & Danhua, W. (2009). Long-term interest rates and expected future budget deficits: evidence from the term structure. *Applied Economics Letter*, Vol. 16, Iss. 4, 365–368.

Von Hagen, J. (2010). Sticking to fiscal plans: the role of institutions. *Public Choice,* Vol. 144 (3–4), 487–503.

Walsh, C. (2010). *Monetary Theory and Policy* (3rd ed.). Cambridge, MA: The MIT Press.

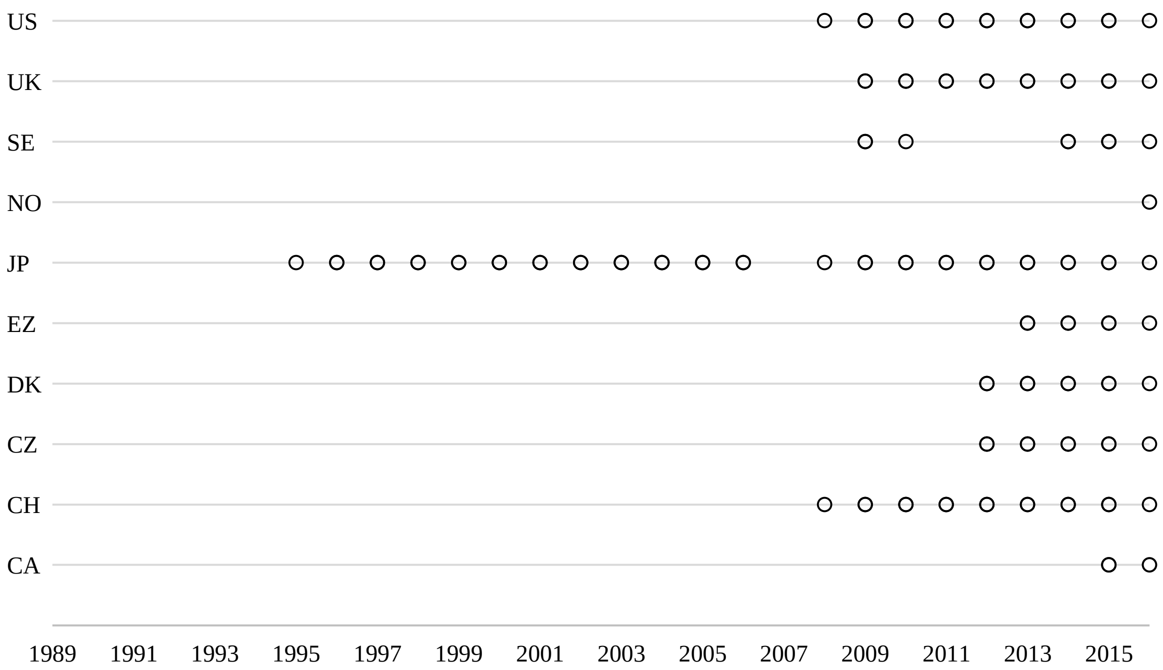
Westerlund, J. (2007). Testing for Error Correction in Panel Data. *Oxford Bulletin of Economics and Statistics*, Vol. 69, Iss. 6, 709–748.

Wooldridge, J. M. (2010). *Econometric Analysis of Cross Section and Panel Data* (2nd ed.). Cambridge, MA: The MIT Press.

Wooldridge, J. M. (2012). *Introductory Econometrics: A Modern Approach* (5th ed.). Mason: South-Western.

Appendix

Appendix A. Countries with central bank policy rates set at 0.5 % or below.



Notes: 'EZ' denotes the Eurozone countries.

Source: Bloomberg and Central bank websites.

Appendix B. Minimum and maximum forecast errors.

		Minimum		Maximum	
Variable	**Issue**	**Value**	**Observation**	**Value**	**Observation**
Short-term interest rate	Jun	-5.59	GR 1996	5.34	IS 2008
	Dec	-5.81	IS 2009	2.86	IT 1992
Inflation	Jun	-7.46	IE 2009	11.94	NO 2000
	Dec	-6.05	IE 2009	9.44	NO 2000
Growth	Jun	-11.47	SK 2009	21.20	IE 2015
	Dec	-9.45	SK 2009	20.09	IE 2015
Primary balance	Jun	-17.90	IE 2010	16.20	IS 2016
	Dec	-18.98	IE 2010	16.60	IS 2015
Public debt	Jun	-41.28	IS 2014	62.75	IS 2009
	Dec	-51.73	IS 2010	43.88	IS 2008

Source: Author's calculations based on data from OECD EO n.45–100.

Appendix C. Credit rating conversions.

Numerical equivalent	Moody’s	S&P	Fitch	
16	Aaa	AAA	AAA	Investment grade
15	Aa1	AA+	AA+	
14	Aa2	AA	AA	
13	Aa3	AA-	AA-	
12	A1	A+	A+	
11	A2	A	A	
10	A3	A-	A-	
9	Baa1	BBB+	BBB+	
8	Baa2	BBB	BBB	
7	Baa3	BBB-	BBB-	
6	Ba1	BB+	BB+	Non-Investment grade
5	Ba2	BB	BB	
4	Ba3	BB-	BB-	
3	B1	B+	B+	
2	B2	B	B	
1	B3	B-	B-	
0	< B3	< B-	< B-	

Source: Author’s conversions based on data from Trading Economics.