Fiscal policy and the business cycle: the impact of government expenditures, public debt, and sovereign risk on macroeconomic fluctuations

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This thesis studies the role of fiscal policy over the business cycle based on a combination of empirical macroeconometric techniques and macroeconomic theory. The focus of the analysis to be conducted is on the impact of government expenditure policies, public debt, and sovereign default risk on short-run to medium-term macroeconomic dynamics. The aim of this analysis is to contribute to fill some of the gaps in contemporary research on fiscal policy that have become obvious in the context of the turbulent economic events over the past few years.

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Fiscal Policy and the Business Cycle

The Impact of Government Expenditures, Public Debt, and Sovereign Risk on Macroeconomic Fluctuations

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aan de Universiteit van Amsterdam
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Amsterdam, July 2011
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List of Abbreviations

a.k.a. also known as
ALFRED Archival Federal Reserve Economic Data
ARRA American Recovery and Reinvestment Act
AWM Area-Wide Model
BEA Bureau of Economic Analysis
BIS Bank of International Settlements
BLS Bureau of Labor Statistics
CBO Congressional Budget Office
CBRT Central Bank of the Republic of Turkey
CE Consensus Economics
cf. confer | compare
CPI Consumer Price Index
DGP data-generating process
DSGE dynamic stochastic general equilibrium
e.g. exempli gratia | for example
EC European Commission
ECB European Central Bank
EMBIG Emerging Market Bond Index Global
EMU Economic and Monetary Union
ESA95 European System of National and Regional Accounts
etc. et cetera | and so forth
EU European Union
G-20 Group of Twenty
GDP gross domestic product
GMM Generalized Method of Moments
HICP Harmonized Index of Consumer Prices
i.e. id est | that is
ILO International Labor Organization
IMF International Monetary Fund
ML maximum likelihood
NIPA National Income and Product Accounts
NOEM new open-economy macroeconomic
OECD Organization for Economic Cooperation and Development
### List of Abbreviations

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<tr>
<td>OLS</td>
<td>ordinary least squares</td>
</tr>
<tr>
<td>PPP</td>
<td>Paredes, Pérez, and Pedregal</td>
</tr>
<tr>
<td>RWM</td>
<td>Random Walk Metropolis</td>
</tr>
<tr>
<td>SPF</td>
<td>Survey of Professional Forecasts</td>
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<tr>
<td>SVAR</td>
<td>structural vector autoregression</td>
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<tr>
<td>TFP</td>
<td>total factor productivity</td>
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<tr>
<td>TVP-VAR</td>
<td>time-varying parameters vector autoregression</td>
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<tr>
<td>U.S.</td>
<td>United States</td>
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<td>UK</td>
<td>United Kingdom</td>
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<tr>
<td>VAR</td>
<td>vector autoregression</td>
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“That these predictions [of Keynesian economics] were wildly incorrect and that the doctrine on which they were based is fundamentally flawed are now simple matters of fact, involving no novelties in economic theory.”


“My advice to the Obama people is to figure out how much help they think the economy needs, then add 50 percent. It’s much better, in a depressed economy, to err on the side of too much stimulus than on the side of too little.”


“The claim that budget deficits make the economy poorer in the long run is based on the belief that government borrowing “crowds out” private investment (...). Under normal circumstances there’s a lot to this argument. But circumstances right now are anything but normal.”


“The time for talk is over. The time for action is now. Because we know if we do not act, (...) [c]risis could turn into catastrophe.”


“For the past month or so, I’ve made it a habit to ask fellow economists how the response to the financial crisis has been improved by the past few decades of macroeconomic theorizing. (...) I don’t have much to report.”

—Justin Wolfers, post on Freakonomics.com, January 28, 2009.
Chapter 1

Introduction

1.1 Macroeconomic background

In the past few years, the world experienced not only the worst financial and economic crisis in decades but also fiscal policy steps of unprecedented scope in response to that crisis. When the financial crisis reached its most critical stage in the autumn of 2008, several governments put in place rescue and support packages for the financial sector. These efforts were strongest in developed countries, which were most affected by the financial turmoil. In 2009, the economic crisis had taken shape and spread across the globe. With monetary policy means exhausted, a large number of governments in developed, developing, and emerging countries introduced fiscal stimulus measures targeting the real economy in an effort to thwart deeper recession.¹

The numbers that describe the fiscal response to the crisis are significant. According to the ILO, 32 countries including all G-20 nations had announced economic stimulus packages by the first quarter of 2009. At that time, the budgets allocated to those packages amounted to about two trillion U.S. dollars or equivalently 1.4% of global GDP. The packages targeting the financial sector were even larger in size, outweighing the stimulus measures by a factor of five or more (see Khatiwada, 2009). According to

¹Detailed timelines of crisis events and the associated international financial sector policies are provided on http://www.newyorkfed.org/research/global_economy/policyresponses.html. Such policies included in particular increased guarantees for private deposits, guarantees for other bank liabilities, capital injections, funds to purchase mortgage bonds and commercial paper, and options to purchase assets of uncertain value. A study of the ILO summarizes the international fiscal stimulus measures that had been announced by the first quarter of 2009, see Khatiwada (2009), and the IMF has provided several updates since then.
the IMF (2009b), the G-20 stimulus planned as of April 2009 was 2% relative to the group’s GDP in 2009 and 1.5% in 2010. Financial sector support, including guarantees, stood at 32% of GDP with upfront financing needs at 3.5%. According to estimates from November 2010, the actual stimulus was 2.1% of GDP in 2009, a planned stimulus of 2% in 2010, and a further stimulus of 1% in 2011 (see IMF, 2010b).

The main goal of the financial sector support measures was to prevent further financial market turmoil or, possibly, a systemic breakdown of the global financial system. Hopes were also raised by several economic advisors that the stimulus measures would be effective in lifting the economy out of recession, particularly in the U.S. where the largest stimulus package was adopted (see, for instance, Romer, 2009, Romer and Bernstein, 2009, Summers, 2008). Such hopes were reinstated by official bodies and policymaking institutions when economic growth started to pick up towards the end of 2009, soon after the implementation of the first stimulus measures. At that time, economic projections indeed indicated that a stimulus-driven recovery was under way. The recovery was expected to be led by the U.S. among the advanced economies and it was forecasted to be even stronger in developing and emerging market economies (see e.g. CBO, 2009; IMF, 2009a, 2010c; OECD, 2009).

However, while the recovery accelerated during 2010 in several emerging market economies, private domestic demand remained weak in the U.S. and in various other advanced economies. In addition, the case for fiscal consolidation became obvious as a sovereign debt crisis erupted in Europe in the spring of 2010 and U.S. public debt was also projected to rise at an accelerated pace (see IMF, 2010e).

This fragile and uneven recovery has continued into 2011, while advanced-country debt sustainability concerns have remained (see IMF, 2011b). At the current and the foreseen fiscal stance, the average public debt-to-GDP ratio in advanced economies is still projected to rise from 73.1 percent in 2007 to 107.3 percent in 2016. Furthermore, despite major policy actions by national governments, EU, ECB, and IMF, the European debt crisis seems far from over as government bond yields and interest rates on credit default swaps are still rising in a number of euro area countries. The dispersion in yields already exceeds pre-EMU levels, including during the European Monetary System crisis of the early 1990s (see IMF, 2011a).
The present thesis studies the relationship between fiscal policy and short-run to medium-term macroeconomic fluctuations (i.e. the business cycle). The events described above have filled the macroeconomic news and the economic and political agenda while I was working on this topic. It has been a turbulent time but, admittedly, also an exciting time to conduct this research.

1.2 Macroeconomic research

What does existing research tell us on the linkages between fiscal policy and the business cycle, with a view of the macroeconomic background described above? This section reviews a number of critical issues, focusing on three relevant topics on which the macroeconomic literature seems to provide relatively little guidance.

1.2.1 Measuring the effects of fiscal policy

First, there is hardly any consensus across empirical studies—using vector autoregression (VAR) techniques or other empirical methods—on the size of fiscal multipliers on economic output, which are often used to measure the effectiveness of discretionary fiscal policy over the cycle. As a by-product of this lack of consensus, there is also significant uncertainty on the size or even the sign of the effects of discretionary fiscal policy on other macroeconomic variables, in particular output components.\(^2\)

The short-run fiscal multiplier on output is defined as the percentage response of GDP in a given period to an *autonomous change* in a given fiscal item or budgetary instrument of size 1% of GDP which occurs in that period.\(^3\) Thus, if a multiplier is found to be larger (smaller) than one, a fiscal expansion tends to crowd in (crowd out) some component or components of private demand. Opinions differ on the definition

\(^2\)Large-scale DSGE models have also been used to estimate the size of fiscal multipliers by Bayesian techniques. A relatively broad consensus on the effectiveness of discretionary fiscal policy has emerged from this literature, see in particular Coenen, Erceg, Freedman, Furceri, Kumhof, Lalonde, Laxton, Lindé, Mourougane, Muir, Mursula, de Resende, Roberts, Roeger, Snudden, Trabandt, and in ’t Veld (2010). I do not provide a complete review here and refer the reader instead to Leeper, Traum, and Walker (2011) for a discussion of the restrictions on fiscal multipliers implied by the commonly used classes of models and the commonly adopted Bayesian priors.

\(^3\)This definition was originally proposed by Kahn (1931) and it is often used in the literature, see e.g. Blanchard and Perotti (2002), Caldara and Kamps (2008), and Ramey (2011a).
of multipliers at longer horizons; some studies relate the response of GDP at a given horizon to the initial change in the relevant fiscal variable (e.g. Blanchard and Perotti, 2002) whereas others use cumulated changes of GDP relative to cumulated changes in fiscal variables up to some horizon, possibly in present value terms (e.g. Mountford and Uhlig, 2009). Given these definitional differences, the focus of the following discussion is on short-run multipliers but the main conclusion (i.e. lack of consensus) is no different as regards the size of longer-term multipliers.

Estimates of fiscal multipliers are indeed dispersed to a degree that there is no agreement across different empirical studies, even across those that use similar techniques, on whether multipliers are usually smaller or larger than one. In particular, according to a recent survey by Ramey (2011a), estimated short-run multipliers for temporary, deficit-financed increases in government purchases of goods and services in the U.S. usually lie between 0.8 and 1.3. However, the data can also not reject 0.5 and 1.8. Ramey (2011a) further notes that there is significant uncertainty on the size of multipliers for fiscal expansions falling on tax cuts. In addition, from a recent survey by Afonso, Baxa, and Slavík (2011) one can conclude that empirical estimates of fiscal multipliers in Europe, even for identical countries, are also quite dispersed. Moreover, the general uncertainty on the effects of fiscal expansions on output goes along with a lack of agreement on the impact of spending expansions on private consumption and investment (see Perotti, 2008).

What are the reasons for this lack of consensus?

One critical issue is that econometric problems in the measurement of structural fiscal shocks (i.e. autonomous changes in fiscal instruments) pose significant challenges to empirical work. An important potential cause is that the presence of news or foresight about future policy changes can create equilibria with non-fundamental or non-invertible moving average representations. This means that structural shocks cannot be recovered by VAR techniques. This issue has been pointed out with reference to fiscal policy by Leeper, Walker, and Yang (2011), Ramey (2011b), and Yang (2005). The issue is especially relevant in the case of fiscal policy due to frequent pre-announcement of fiscal measures and legislative lags or other delays in the implementation of announced measures. The above authors show that non-fundamentalness
due to policy foresight can seriously distort VAR inference of the effects of fiscal measures. If the issue is ignored, estimated multipliers can even have opposite signs than implied by an underlying model that does incorporate policy foresight.

A second difficulty is presented by the fact that the causal effects of fiscal policy are still hard to identify even in the absence of policy foresight, or also in the presence of policy foresight when the above non-fundamentalness issues could be circumvented in some way. The main reason is that both government expenditures and revenues, to some extent, automatically respond to economic fluctuations. Such changes need to be distinguished through appropriate identification approaches from deliberate policy changes. If the latter is not accomplished, such endogenous reactions of fiscal variables to the business cycle can induce reverse causation. This can lead to biased and therefore incorrect estimates of the effects of fiscal policy.\footnote{See Caldara and Kamps (2008) for an overview and a comparison of differences in outcomes of existing methods for the identification of fiscal shocks.}

A third problem are instabilities over different time periods, whose existence has been pointed out by several studies, including e.g. Bénassy-Quéré and Cimadomo (2006), Bilbiie, Meier, and Müller (2008), and Blanchard and Perotti (2002). Possible factors of instability include structural changes and breaks that could lead to changes in the effects of fiscal policy over time. For example, increasing trade integration could lead to increasing open-economy leakages of fiscal expansions. For obvious reasons, such types of sub-sample instabilities make it hard to interpret estimation results for overlapping time periods without ambiguity.

Hence, it is hard to measure the effects of fiscal policy on output and other variables, which may explain part of the missing consensus in the empirical literature. The aim of Chapters 2 and 3 is to address some of the issues just discussed.

\subsection*{1.2.2 Fiscal policy during financial crises}

A second aspect are the effects of fiscal policy in times of financial stress, on which relatively little is known on both the theoretical and empirical side. The lack of theoretical studies has been associated with a neglect of relevant linkages between the real economy and the financial sector in standard macroeconomic models. On the other
hand, some of the lack of empirical studies might be explained by a relative shortage of data on crises in advanced economies.\footnote{Noteworthy developed-country financial and economic crisis episodes, according to Spilimbergo, Symansky, Blanchard, and Cottarelli (2009), include the U.S. Savings and Loans crisis in the 1980s, the Nordic countries in the early 1990s, Japan in the 1990s, and Korea in 1997.} An overall scarcity of data for developing and emerging market economies, where crises have been more frequent, and concerns on the quality of the available data contribute to the lack of relevant studies.

Having said that, the empirical literature on the topic is slowly growing. For example, Afonso, Baxa, and Slavík (2011) apply a threshold VAR approach using quarterly data for the U.S., the UK, Germany, and Italy to investigate whether the effects of fiscal policy on economic activity differ depending on financial market conditions. The authors conclude that there are only small differences in the effects of fiscal shocks in regimes of high financial stress compared to regimes of low financial stress.

Cross-sectional studies include, for instance, Baldacci, Gupta, and Mulas-Granados (2008) who use OLS and ordered logit to estimate the effects of fiscal policy interventions during 118 episodes of banking crises in developed and emerging countries. These authors find that fiscal stimulus accompanied by financial sector policies can shorten such crises, but this result does not hold for countries with limited fiscal space. On the other hand, for a panel of 127 OECD and non-OECD countries, Afonso, Grüner, and Kolerus (2010) cannot reject the hypothesis that the effects of fiscal policy are the same in normal times and during a financial crisis.

Hence, the evidence from recent empirical studies does not yet speak very clearly on the effects of fiscal policy during financial crises. Of course, similar problems as in the measurement of fiscal multipliers also affect those studies. In addition, the available data is often not rich enough to distinguish between different transmission channels (e.g. exceptional financing constraints) or different policy instruments. It therefore seems important to put more theoretical work into analyzing the effects of fiscal policy in the presence of financial frictions. Structural macroeconomic models can be used to conduct this type of analysis, but standard models need to be augmented by adding the relevant macroeconomic relations and frictions.

In the face of the recent crisis, there has been significant progress at this frontier. A relative large literature has developed that studies the interaction of fiscal and mon-
etary policy in a crisis scenario, analyzing the effects of fiscal stimulus in structural macroeconomic models when monetary policy allows real interest rates to fall or when nominal interest rates are at the zero lower bound, as occurred during the recent crisis (see e.g. Christiano, Eichenbaum, and Rebelo, 2009; Coenen et al., 2010; Davig and Leeper, 2011; Eggertsson, 2011; Erceg and Lindé, 2010; Woodford, 2011).

When attempting to study potential financial sector feedbacks of such policies, one faces the problem that standard models are not set up for an analysis of this type. However, promising models with frictions in financial intermediation have recently been developed. These models have been used to study the effects of central bank credit intermediation and government policies targeting the financial sector in a financial crisis; see, in particular, Gertler and Karadi (2011), Gertler and Kiyotaki (2010), Gertler, Kiyotaki, and Queralto (2010), and also Christiano and Ikeda (2011).

However, many questions remain open. For instance, what kind of interactions can we expect due to the presence of government securities holdings next to private assets on bank balance sheets? What are the effects of traditional discretionary government policies (such as spending expansions) and policies targeting the financial sector if these are financed by issuing bonds to a troubled banking sector? In particular, can higher government deficits affect bank lending to the non-financial sector? These types of questions have played an important role in recent policy discussions (see e.g. IMF, 2010a), but standard macroeconomic models are not yet able to guide such discussions. The aim of Chapter 4 is to tackle this issue.

1.2.3 Sovereign risk and macroeconomic fluctuations

The literature furthermore tends to lack quantitative business cycle models that take into account the possibility that governments can default on their debt. This deficiency became obvious as post-crisis fiscal sustainability concerns have recently come to the forefront of the economic and political agenda in the developed world.

In particular, as argued by Bi and Leeper (2010), policy evaluations in models that do not allow for the possibility of sovereign debt default seem unreliable when applied to economies where financial markets regard government debt as risky. It would therefore be useful to characterize and understand the link between public debt,
sovereign risk, and macroeconomic fluctuations to be able to recommend appropriate fiscal or monetary policies in an environment where debt sustainability is a concern. To address this issue, it would be useful to estimate a business cycle model on a sample of macroeconomic data that includes at least some episode where sovereign default risk has played a significant role. However, the data for advanced economies does not seem very informative in this respect given the lack of applicable episodes over the past few decades. At the same time, the literature on quantitative business cycle models for emerging market economies, where such episodes have again been more frequent than in advanced economies, is still relatively small.

Some studies have however analyzed emerging market business cycles in calibrated or estimated business cycle models. In particular, Aguiar and Gopinath (2007) examine the role of permanent productivity shocks. In addition, Chang and Fernández (2010), García-Cicco, Pancrazi, and Uribe (2010), Neumeyer and Perri (2005), and Uribe and Yue (2006) explore the impact of financial frictions such as debt-elastic interest rates. One broad conclusion that emerges from this literature is that unlike permanent productivity shocks, financial frictions can explain important regularities of emerging market business cycles, in particular the relatively high observed volatility of consumption relative to output and the countercyclicality of interest rates. Given this evidence and the relative frequency of default episodes in emerging countries, it seems promising to continue in this direction by focusing on financial frictions that are explicitly linked to the risk of sovereign debt default.

Theoretical analyses of sovereign default risk include, in particular, Schabert (2010), Schabert and van Wijnbergen (2011), and Uribe (2005). A common conclusion of these studies is that the possibility of sovereign default matters for the implementation of monetary policy. A quantitative analysis of sovereign default risk is provided by Juessen, Linnemann, and Schabert (2009), who develop a real business cycle model that allows for government debt default when fiscal policy does not preclude a Ponzi game. This study shows that default premia can emerge at relatively high debt-to-GDP ratios. In addition, Bi and Leeper (2010) develop a real business cycle model

\[ \text{Related work studies “fiscal limits” to debt accumulation and the interactions of fiscal and monetary policy in the absence of nominal sovereign debt default; see, in particular, Leeper and Walker (2011), and Davig, Leeper, and Walker (2010).} \]
that allows for debt default when the government reaches the limit of its capacity to raise revenues through distortionary taxation. These authors show that certain types of fiscal reforms can shift this limit to prevent default premia from emerging. Moreover, a recent study by Mendoza and Yue (2011) makes a link between quantitative models on strategic sovereign default based on Eaton and Gersovitz (1981), where default events are driven by exogenous output endowment processes, and quantitative models of emerging market business cycles with debt-elastic interest rates.\(^7\) This study shows that a model that jointly determines the equilibrium dynamics of output and sovereign default does well in explaining key stylized facts of actual defaults.

In spite of this progress, to my knowledge no study has attempted to analyze the implications of sovereign default risk in a quantitative structural model estimated by full information methods. This type of analysis would however allow to describe the joint behavior of different economic time series conditional on a system of structural macroeconomic relations, potentially providing additional understanding of macroeconomic amplification and propagation channels due to sovereign risk. The aim of Chapter 5 is to make progress on this matter.

**1.3 Overview of the thesis**

Overall, despite significant recent progress, the above discussion indicates that the macroeconomic literature

(i) has not yet reached consensus on the size of fiscal multipliers due to, in particular, problems in measuring fiscal shocks and, moreover, sub-sample instability,

(ii) provides only relatively little guidance on appropriate fiscal policies in a situation of financial stress, and

(iii) shows scope for further research on quantitative business cycle models that incorporate sovereign default risk.

The aim of this thesis is to provide a contribution in filling those gaps. Using a combination of empirical macroeconometric techniques and macroeconomic theory, the thesis studies the impact of government expenditure policies, public debt, and

\(^7\)Recent studies on strategic default include, for instance, Aguiar and Gopinath (2006), Arellano (2008), and Yue (2010).
sovereign default risk on business cycle fluctuations. In the following chapters, the thesis addresses the three issues listed above. Developments in public debt are of course linked to all three issues. However, for reasons that are discussed below, the thesis puts emphasis on expenditure policies whereas measures falling on the revenue side of the government budget (i.e. tax policies) are not analyzed.

Chapter 2, which is based on joint work with Jacopo Cimadomo and Sebastian Hauptmeier, provides an analysis of time variation in the macroeconomic effects of government consumption and investment spending. As mentioned above, empirical studies of the effects of discretionary fiscal policy usually do not take into account the possibility that those effects could change over time. In most studies, the effects of fiscal policy are instead estimated on average over samples spanning around two decades or more. However, presuming that the structure of an economy can change during such a period, it seems cautious not to exclude the possibility that fiscal policy could have different effects at different points of time.

The chapter thus estimates VAR models with time-varying parameters for the euro area. This particular method is chosen as it allows for a flexible description of time variation in the relationship among macroeconomic variables. The chapter then identifies structural shocks to government spending at different points of time and simulates the short-run to medium-term effects of those shocks. The chapter also describes potential sources of the detected time variation using simple regression analysis. The latter is thought to add additional structure to the results and thereby to contribute to the understanding of the fiscal transmission mechanism. The focus on the euro area is motivated by the facts that the empirical literature is especially inconclusive on the effects of fiscal policy in Europe and that sub-sample instability due to structural changes is an obvious possibility at the euro area level. The focus on government spending stems from the advantage that significant endogenous reactions to macroeconomic fluctuations seem less likely in the case of public expenditures than in the case of tax revenues, making it easier to identify autonomous policy changes.

Chapter 3 focuses on the econometric problems that are posed to structural VAR analysis by the presence of news or foresight on fiscal policy, following in particular the analysis of Leeper, Walker, and Yang (2011). As argued above, policy foresight
presents significant challenges to empirical work since it may lead to non-invertibility of the moving average representations of the relevant equilibrium time series into VAR representations. However, Laubach (2008) has already pointed out the possibility of using direct measures of expectations on fiscal variables, such as survey data, in order to address those challenges. Several recent contributions have indeed used information from forward-looking data to tackle the econometric issues due to policy foresight. The chapter seeks to provide a theoretical foundation for such attempts.

The chapter is concerned with the particular problem of quantifying the effects of government spending under policy foresight. Based on a simple theoretical model, the chapter first shows how the associated econometric issues can be addressed by using data that captures the expectations of economic agents (or market participants) on future government spending in VAR models, and how such an approach makes it again possible to identify structural spending shocks by VAR methods. The chapter then estimates the effects of government spending in the U.S., using data from the Survey of Professional Forecasters to measure the relevant expectations. The renewed focus on government spending also in this chapter, in particular the sum of government consumption and government investment in the U.S., is due to the fact that quarterly survey data is only available for this sum. Similar data is however not available for other budgetary items, for the U.S. or other countries.

Chapter 4, which is based on joint work with Sweder van Wijnbergen, takes a step towards an analysis of government policies in an environment of financial stress. This chapter builds on recent work by Gertler and Karadi (2011) who have developed a New Keynesian structural macroeconomic model with financial frictions due to an agency problem in financial intermediation. The particular type of friction proposed by Gertler and Karadi—also used in Gertler and Kiyotaki (2010)—leads to endogenous balance sheet constraints on the operations of financial intermediaries. These constraints imply a financial accelerator mechanism that helps to generate key features of a financial and economic crisis of the type and the magnitude of, not exclusively, the recent crisis. Those features include, in particular, mutual feedbacks between financial sector balance sheets and the real economy. However, the above studies assume that the government does not rely upon intermediary funding. As argued in the chapter, this assumption
does not do justice to the actual practice of fiscal financing.

The chapter therefore extends the above framework by allowing for the presence of government securities in intermediary portfolios. This extension makes it possible to analyze the effects of government policies during a financial crisis when such policies are financed at least to some extent through the relevant financial intermediaries. The chapter then analyzes the effects of deficit-financed stimulus measures and financial sector policies. The particular set of policies that is used suitably captures the main fiscal policy measures that were applied during the recent crisis. The chapter investigates how the presence of intermediary balance sheet constraints in interaction with portfolio adjustments can affect the effectiveness of those policies.

Chapter 5, which is based on joint work with Malte Rieth, analyzes the role of sovereign default risk as a driving factor of macroeconomic fluctuations. The analysis is based on Schabert and van Wijnbergen (2011), who set up a New Keynesian small open economy model that takes into account the possibility that a conventional fiscal rule with a feedback from higher debt levels on taxes can imply politically infeasible rates of taxation. In that case, the government defaults on (part of) its outstanding debt. Since investors rationalize this possibility, the latter leads to default premia that affect the expected return on government bonds and that are endogenously linked to the stock of real government liabilities. This model thus describes an environment where the possibility of sovereign debt default is a relevant concern.

The chapter extends this model by allowing for foreign currency denominated debt to reflect the typical situation in emerging market economies, where governments can usually not borrow in their own currency abroad (a.k.a. the “original sin” problem, Eichengreen and Hausmann, 1999). The model is then estimated by Bayesian full-information techniques on data for an emerging market economy, taking Turkey’s experience as a natural experiment. In particular, Turkey was hit by a severe financial crisis in November 2000 when nominal interest rates increased sharply, accompanied by a downgrading of government debt to below investment grade. The Turkish experience therefore reflects a situation where fears of sovereign debt default have played an important role, although a debt default did not actually occur. Based on the estimated model, the chapter assesses the role of sovereign default risk in explaining business
cycle fluctuations in this type of emerging market environment.

Chapter 6 summarizes the results of the analysis in Chapters 2 to 5 and provides an overall conclusion that emerges from those results. This final chapter also qualifies the results in the light of what has been done and what has not been done in this thesis. The chapter ends with a suggestion for future research.
Chapter 2

Transmission of Government Spending Shocks in the Euro Area: Time Variation and Driving Forces*

Abstract

This chapter applies structural vector autoregressions with time-varying parameters to investigate changes in the effects of government spending shocks in the euro area and it studies the driving forces of those changes. We first present evidence on the effects of government spending shocks on real GDP and other variables for individual quarters during the period 1980-2008. We then exploit state dependency using regression inference to add additional structure to the results. Our findings show that short-run spending multipliers have increased from the early 1980s to the late 1980s but they have decreased thereafter until the late 2000s. Moreover, the longer-term effects of spending shocks have declined substantially over this period. We also find that the time variation in spending multipliers can be traced back to increasing availability of credit and rising debt-to-GDP ratios, as well as a smaller share of government investment and a larger share of public wages in total spending.

2.1 Introduction

Fiscal policy has been rediscovered as a tool for short-run economic stabilization. In the context of the recent financial and economic crisis, governments around the world have enacted unprecedented fiscal stimulus packages to counter the severe economic downturn. For instance, the fiscal stimulus adopted within the European Economic

*This chapter is based on joint work with Jacopo Cimadomo and Sebastian Hauptmeier.
Recovery Plan is expected to reach about 1% of the EU’s GDP in 2009 and 0.9% in 2010, and it is largely based on government spending (see European Commission, 2009). However, there is a high degree of uncertainty concerning the macroeconomic impact of government expenditure policies. The theoretical and empirical literature on the effects of government spending shocks reflects this uncertainty as it is rather inconclusive so far, especially as regards the euro area.

Against this background, this chapter offers two contributions. First, we uncover changes in the effects of government spending shocks in the euro area over the period 1980-2008 using the tools of time-varying parameters VAR (TVP-VAR) analysis, allowing for drifting coefficients and stochastic volatility in the VAR model. Second, we provide empirical evidence on the driving forces of the detected time variation in spending multipliers. In particular, using regression inference we relate the estimated multipliers to a set of macroeconomic indicators and to the composition of spending. The underlying idea is that this type of analysis can add additional structure to the results, in a way that may reveal useful information on the fiscal transmission mechanism. To our knowledge, this is the first study that investigates time variation in the effects of government spending shocks through an application of TVP-VAR techniques.¹ In addition, the present study represents the first attempt to provide empirical evidence, by means of a systematic exploitation of state dependency, on the driving factors behind the changing patterns of spending multipliers.

Our analysis is based on a quarterly fiscal data set for the euro area developed by Paredes, Pedregal, and Pérez (2009) for the period 1980-2008. The focus on the aggregate euro area has several advantages. In particular, sub-sample instability has been an obvious possibility at the euro area level, given significant structural changes experienced since the 1980s. Examples include the adoption of the Maastricht Treaty in 1992, the run-up to the EMU, the introduction of the single currency, and the single monetary policy since 1999. Such events should enhance the scope for time variation and help the identification of the key elements of the fiscal transmission mechanism.

¹TVP-VAR models have been applied to study changes in the effects of monetary policy and the relation to the “Great Moderation” (see e.g. Benati and Muntaz, 2007; Canova and Gambetti, 2009; Cogley, Primiceri, and Sargent, 2010; Cogley and Sargent, 2002, 2005; Galí and Gambetti, 2009; Primiceri, 2005), and the implications of structural change for macroeconomic forecasts (see D’Agostino, Gambetti, and Giannone, 2009).
In addition, while an investigation of time variation at the country level would also be of interest, such an analysis would suffer from the lack of fiscal data sets for single euro area countries of sufficient quality and length.\(^2\)

On the other hand, the use of aggregate euro area data also poses the question of how fiscal shocks should be interpreted in this context. In particular, while the single monetary policy has been in place since 1999 and national monetary policies were largely synchronized before that date, fiscal policy, despite a higher degree of coordination within the EU fiscal framework, remains mainly a country-specific matter. At the same time, the use of aggregate data is justified for the following reasons. First, there is evidence that discretionary fiscal policies have co-moved significantly over the past two decades at the EU and euro area level (see e.g. Giuliodori and Beetsma, 2008). Second, aggregate euro area data can be interpreted as a weighted average of the corresponding country-specific components. This interpretation does not necessarily require that national fiscal policies are aligned. What is instead required is that a spending shock—country-specific or coordinated—is large enough to have an identifiable impact on euro area aggregates. Results are then likely to be driven by shocks occurring in those countries which have the largest weight in euro area variables. Empirical support for this view is provided by Bruneau and Bandt (2003), who show that euro area fiscal shocks were largely induced by Germany, especially in the 1990s. Against this background, a growing number of studies, based on calibrated or estimated DSGE models, now postulates an aggregate fiscal policy for the euro area.\(^3\)

Based on a fixed parameters VAR model estimated over the 1980-2008 sample, our first set of results indicates that, on average, government spending shocks have had an expansionary short-run impact and moderately contractionary longer-term effects on output and the components of domestic private demand in the euro area. However, our time-varying approach uncovers important changes in the effects of spending shocks. In particular, our results suggest that short-run spending multipliers have increased between the early 1980s and the late 1980s but they have decreased thereafter. More- 

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\(^2\)While the dataset provided by Paredes et al. (2009) is consistent with official national accounts data according to European System of Accounts standards (ESA95), this is not the case for (quarterly) fiscal data of single euro area countries, at least for the period preceding 1999.

\(^3\)See e.g. Christoffel, Coenen, and Warne (2008), Fagan, Henry, and Mestre (2005), Forni, Monteforte, and Sessa (2009), Ratto, Roeger, and in ’t Veld (2009), and Smets and Wouters (2003).
over, the expansionary effects of government spending have become more short-lived over time, and the estimated longer-term effects have decreased substantially. The effects of spending-based fiscal expansions on output indeed appear to be particularly low in the current decade. In addition, smaller spending multipliers on output are found to coincide with a weaker response of private consumption and the real wage, but a stronger response of the short-term nominal interest rate.

With respect to the driving forces of the detected time variation, our evidence points towards availability of household credit as an important determinant of the size of short-run spending multipliers. This result underpins arguments suggesting that access to credit or non-Ricardian behavior by households matter for the effectiveness of fiscal expansions. We also find that a smaller share of investment expenditures and a larger wage component in total government spending are associated with smaller short-run multipliers. Our results therefore seem to provide support for the view that government investment can have positive aggregate supply effects in addition to the aggregate demand effect of government purchases. The fact that wage payments are associated with lower multipliers supports arguments stating that government wage expenditures may have adverse effects in an imperfect labor market through their impact on reservation wages (see Alesina and Ardagna, 2010). Finally, we find that higher debt-to-GDP ratios are associated with lower spending multipliers at longer horizons. This result might suggest that, given higher financing needs of euro area governments, sustained deficits after a spending shock could lead, for instance, to rising concerns on the sustainability of public finances or expectations of a larger future consolidation, which according to our results seems to depress private demand.

The remainder of the chapter is organized as follows. Section 2.2 reviews the related literature. Section 2.3 describes our econometric models, the estimation method, the data, and the structural identification approach. Section 2.4 presents estimation results for the identified models. Section 2.5 discusses several robustness checks. Section 2.6 investigates the driving forces of the detected time variation. It first discusses theoretical views on the fiscal transmission mechanism and, based on this discussion, it identifies the determinants underlying the time variation in spending multipliers using regression analysis. Section 2.7 concludes.
2.2 Related literature

On the theoretical side, there is still considerable disagreement concerning the impact of government spending shocks on important macroeconomic variables. Macroeconomic models used to evaluate the effects of fiscal policy tend to diverge in their predictions (cf. Cogan, Cwik, Taylor, and Wieland, 2010). Neoclassical models with optimizing agents and flexible prices typically indicate a rise in output but a fall in private consumption and real wages following an exogenous increase in government goods purchases (see e.g. Baxter and King, 1993). New Keynesian models, on the other hand, can generate an increase in real wages depending on the monetary regime (see Linnemann and Schabert, 2003). However, basic versions of those models also tend to predict a crowding out of private consumption unless additional features are included which dampen the negative wealth effect of a fiscal expansion. Examples include non-Ricardian consumers (Galí, López-Salido, and Vallés, 2007), imperfect substitutability between public and private consumption (Linnemann and Schabert, 2004), small wealth effects on labor supply (Monacelli and Perotti, 2008), and spending expansions followed by reversals, which create expectations on a future fall in real interest rates (Corsetti, Meier, and Müller, 2009; Corsetti, Kuester, Meier, and Müller, 2010).

On the empirical side, the effects of government spending shocks are typically investigated within the structural VAR (SVAR) framework. Alternatives include the event-study approach of Ramey and Shapiro (1998) or, more recently, Ramey (2011b). Despite an increasing number of studies in this field, many questions remain open. In particular, the effects of government spending shocks in the euro area are largely unexplored. The scarcity of empirical results for the euro area as a whole and also for euro area countries has been mainly due to the lack of quarterly fiscal data, a limitation which has been overcome recently through a quarterly fiscal database for the euro area compiled by Paredes et al. (2009). This data set, which covers the period 1980Q1-2008Q4, is coherent with official annual and quarterly national accounts data, as far as

\footnote{See e.g. Blanchard and Perotti (2002), Caldara and Kamps (2008), Fatás and Mihov (2001), Mountford and Uhlig (2009), and Perotti (2005).}

\footnote{Ramey and Shapiro (1998) and Ramey (2011b) are concerned with the possibility that autonomous fiscal policy changes might be anticipated in advance of their implementation, which is an important challenge for the validity of SVAR results. In this chapter, this issue is addressed in Section 2.5, where we discuss several exercises related to the possible anticipation of the identified SVAR shocks.}
quarterly fiscal data is available from national accounts (mostly for the period 1999Q1 onwards). Based on this data set, Burriel, de Castro, Garrote, Gordo, Paredes, and Pérez (2010) show that the qualitative responses of macroeconomic variables to fiscal shocks in a (weighted) representative euro area country compare well with results for the U.S. and previous results for some EU countries.

There is also disagreement on whether the effectiveness of fiscal policy has changed over time, and if so to what extent and why. This lack of disagreement concerns especially the effects of government spending, as the literature lacks empirical tests of possible explanations for changing effects of spending shocks. Blanchard and Perotti (2002) have already emphasized that the size of spending multipliers on output in the U.S. varies considerably across sub-periods. Similarly, based on sub-sample or rolling-windows estimation, Bénassy-Quéré and Cimadomo (2006), Bilbiie et al. (2008), Caldara and Kamps (2008), and Perotti (2005) conclude that the responses of the U.S. and of some European economies to fiscal policy shocks have become weaker in the post-1980 period. Perotti (2005) argues that relaxation of credit constraints, a stronger real interest rate response, and changes in monetary policy could explain the decline in the effects of government spending on GDP and its components in OECD countries. Using a New Keynesian model, Bilbiie et al. (2008) show that the more active monetary policy in the Volcker-Greenspan period and increased asset market participation can explain lower spending multipliers in the U.S. after 1980. Overall, confronting potential explanations for changes in the effects of government spending shocks with additional empirical evidence seems a useful contribution to this literature.

2.3 Econometric methodology

Our empirical approach is based on Bayesian estimation techniques. We prefer a Bayesian approach over estimation by classical statistical methods for reasons discussed by Primiceri (2005). Most importantly, a Bayesian approach facilitates the estimation of time variation in multivariate models with drifting coefficients and stochastic volatility. The main advantage of Bayesian techniques is related to the high dimensionality and the non-linearity of such an estimation problem. By using prior information and by
splitting up the original problem into a number of smaller steps, Bayesian methods are able to deal with the high dimension of the parameter space and possible non-linearities in the likelihood function associated with the estimation problem.

We also prefer the TVP-VAR methodology to simpler methods including sub-sample or rolling-windows estimation for the following reasons. First, structural changes could take the form of long-lasting processes, which would not be reflected in an optimal way by sub-sample estimation; they could come suddenly, which would not be reflected by rolling-windows estimation; they could also come suddenly and be reversed afterwards, which would not be reflected in this way by either type of method. Second, structural changes might not be easily identified a priori. Third, one can think of various alternative structural changes which might impact on the effectiveness of fiscal policy, e.g. monetary policy regime changes or trade integration. It would therefore be difficult to date breaks and to determine the size of rolling windows. The TVP-VAR methodology allows to address these issues through estimates for individual quarters.

2.3.1 Reduced-form VAR models

We consider two versions of a reduced-form VAR of lag order $p$. The first version has fixed parameters:

$$y_t = B_1 y_{t-1} + \cdots + B_p y_{t-p} + \Gamma z_t + u_t, \quad t = 1, 2, 3 \ldots, T, \quad (2.1)$$

where the vector $y_t$ includes government spending, output, private consumption, the short-term interest rate and possibly other macroeconomic indicators. The $B_i$, $i = 1, 2, 3, \ldots, p$, are matrices of coefficients. The vector $z_t$ collects exogenous variables with parameter loadings $\Gamma$. The vector of innovations $u_t$ is assumed to be Gaussian white noise with mean zero and covariance matrix $R$, i.e. $u_t \sim N(0, R)$.

The second version generalizes (2.1) by allowing for drifting coefficients and stochastic volatility in the innovations.\footnote{Our specification of the TVP-VAR follows Cogley and Sargent (2002, 2005) and Primiceri (2005). We apply some additional restrictions on the hyperparameters as discussed below.} Both aspects are supposed to capture structural changes such as shifts in private sector behavior and/or changes in the conduct of pol-
icy. Drifting coefficients are thought to capture changes in the propagation of shocks throughout the economy. Stochastic volatility is introduced to allow for changes in the distribution of the shocks. Hence:

\[ y_t = B_{1,t}y_{t-1} + \cdots + B_{p,t}y_{t-p} + \Gamma_t z_t + u_t, \quad t = 1, 2, 3 \ldots, T, \tag{2.2} \]

where \( u_t \sim N(0, R_t). \) Stack the VAR coefficients by equations in a vector \( \beta_t = \text{vec}(F_t'), \) where \( F_t = [B_{1,t}, \ldots, B_{p,t}, \Gamma_t] \) and \( \text{vec}(\cdot) \) is the column stacking operator. This state vector of coefficients is assumed to follow a driftless random walk:

\[ \beta_t = \beta_{t-1} + \varepsilon_t, \tag{2.3} \]

where \( \varepsilon_t \sim N(0, Q). \) Further, the innovation covariance matrix can be decomposed using a triangular factorization of the form

\[ R_t = A_t^{-1}H_t(A_t^{-1})', \tag{2.4} \]

where \( A_t^{-1} \) is lower triangular with ones on the main diagonal and \( H_t \) is diagonal. Stack the elements below the main diagonal of \( A_t \) row-wise in a vector \( \alpha_t. \) Collect the diagonal elements of \( H_t \) in a vector \( h_t. \) Similarly as the coefficient states, the covariance and volatility states are modeled as (geometric) random walks:

\[
\begin{align*}
\alpha_t &= \alpha_{t-1} + \nu_t, \\
\log h_t &= \log h_{t-1} + \omega_t,
\end{align*}
\tag{2.5}
\]

where \( \nu_t \sim N(0, S) \) and \( \omega_t \sim N(0, W). \) Thus, following Primiceri (2005), both the diagonal elements and off-diagonal elements of the reduced-form covariance matrix can...
drift over time, where the latter allows for changes in the contemporaneous relations among the endogenous variables.

The joint distribution of shocks is postulated as \([u_t, \varepsilon_t, \nu_t, \omega_t]' \sim N(0, V_t)\), where \(V_t\) is block diagonal with blocks \(R_t, Q, S\) and \(W\). Notice that an unrestricted covariance matrix would drastically increase the number of parameters and thus complicate the estimation problem. Independence of \(R_t\) and the hyperparameters implies that the innovations to the VAR parameters are uncorrelated with the VAR residuals. This assumption seems plausible since the innovations capture business cycle events, policy shocks, or measurement errors. It seems unlikely that such short-term events are related to longer-term institutional changes and other changes in the structure of the economy, which are captured by the innovations to the VAR parameters. For example, it can be argued that the introduction of the single currency in the euro area has not been related to technology shocks, government spending shocks, and so on.

We make the additional assumption that \(Q, S,\) and \(W\) are diagonal to further reduce the dimensionality of the problem and to simplify inference. The assumption of (block) diagonality of \(S\) ensures that the rows of \(A_t\) evolve independently such that the covariance states can be estimated row by row (cf. Primiceri, 2005). Diagonality of \(W\) implies that the volatility states are independent such that the simple univariate algorithm of Jacquier, Polson, and Rossi (1994) can be applied to each element of \(u_t\) in order to estimate the volatility states. The reduction of estimated parameters resulting from the diagonality restrictions on \(Q\) and \(S\) helps to save degrees of freedom in our relatively short euro area data set.

### 2.3.2 Estimation method

Both VAR models described above are estimated by Bayesian methods. For the version with fixed parameters, our prior and posterior for the coefficient matrices \(B_i, i = 1, \ldots, p, \Gamma,\) and the covariance matrix \(R\) belong to the Normal-Wishart family with a diffuse prior centered on OLS estimates over the full sample. For the TVP-VAR, we apply a variant of the Gibbs sampler (see Geman and Geman, 1984; Smith and Roberts,
The main steps of the estimation algorithm are outlined here whereas Appendix 2.A provides a detailed description. The Gibbs sampler iterates on the following four steps, sampling in each step from lower dimensional conditional posteriors as opposed to the joint posterior of the whole parameter set.

(i) **VAR coefficients.** Conditional on the data and a history of covariance and volatility states, the observation equation (2.2) is linear with Gaussian innovations and a known covariance matrix. The coefficient states $\beta_t$ can thus be sampled using the Kalman filter and a backward recursion, as described in Carter and Kohn (1994) and Cogley and Sargent (2002).

(ii) **Elements of $A_t$.** Conditional on the data and a history of coefficient and volatility states, equation (2.2) can be rewritten as $A_t u_t = v_t$, with $\text{cov}(v_t) = H_t$. This is a linear Gaussian state space system with independent equations, due to the (block) diagonal structure of $S$ (see Primiceri, 2005). The algorithm of Carter and Kohn (1994) can thus be applied equation by equation to sample the elements of $A_t$ on each row below the main diagonal.

(iii) **Elements of $H_t$.** Conditional on the data and a history of coefficient and covariance states, the orthogonalized innovations $v_t$ are observable. Given the diagonal structure of $W$, the diagonal elements of $H_t$ can be sampled using the univariate algorithm of Jacquier, Polson, and Rossi (1994) element by element, following Cogley and Sargent (2005).

(iv) **Hyperparameters.** Conditional on the data and the parameter states, the state innovations $\varepsilon_t$, $\nu_t$, and $\omega_t$ are observable. This allows to draw the hyperparameters (i.e. the elements of $Q$, $S$, and $W$) from their respective distributions.

Under relatively weak regularity conditions (see Roberts and Smith, 1994) and after convergence, iterations on these steps produce a realization from the joint posterior distribution. We generate 60,000 draws from the Gibbs sampler, of which we burn the first 50,000 to let the Markov chain converge to its ergodic distribution. Of the remaining 10,000 draws, we keep every 10th draw to break the autocorrelation of

---

This leaves us with 1,000 draws from the joint posterior distribution of the model parameters. Appendix 2.C analyzes the convergence properties of the Markov chain, concluding that these properties are overall satisfactory.

We follow conventional choices in the calibration of the priors, similar as in Primiceri (2005), but we take a somewhat more conservative stance on the degrees of freedom of the prior distributions which we set to the minimum value allowed for the priors to be proper. Appendix 2.B provides details on the calibration of the priors while the robustness of the results to alternative choices is analyzed in Section 2.5. Unlike most previous studies, we do not truncate the posterior distribution of the VAR coefficients by discarding draws which do not satisfy stationarity conditions. Cogley and Sargent (2002, 2005) have proposed such a restriction for U.S. monetary policy, arguing that the Fed had ruled out unstable paths of inflation. A similar point is harder to defend for aggregate euro area fiscal data since fiscal variables may have followed unstable paths in some countries. We do however check the robustness of the results by imposing stationarity conditions in Section 2.5.

### 2.3.3 Data description

Our benchmark VAR specification includes government spending, defined as government consumption plus government investment following most of the literature, GDP, private consumption (all in real per capita terms), and the short-term nominal interest rate for the euro area over the period 1980Q1-2008Q4. Real GDP measures economic activity. Private consumption is included since it is the largest component of aggregate demand, and also to be able to contribute to the ongoing discussion on the effects of government spending shocks on that variable. The short-term interest rate is added to this small-scale VAR to assess the impact of government spending shocks on interest rates, and potential changes thereof. We also examine the impact of spending shocks on a broader set of macroeconomic indicators including private investment, net taxes.

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10 The Gibbs sampler is a dependence chain algorithm. However, independent draws should be used when calculating statistics of interest such as posterior means and impulse responses.

11 Perotti (2005) argues that the long-term interest rate has a closer relation to private consumption and investment decisions than the short-term interest rate. Replacing the short-term interest rate by the long-term interest rate did however not lead to any significant changes in our results.
Figure 2.1: Data used in the benchmark specification

![Graphs of Gov. Spending over GDP, Private Consumption over GDP, and Short-Term Interest Rate over time]

Notes. Euro area data, 1980Q1-2008Q4; gov. spending equals final general gov. consumption plus gov. investment; gov. spending and private consumption are expressed as nominal shares of GDP; the short-term interest rate is measured in nominal, annual terms; source of fiscal data: Paredes et al. (2009); source of remaining data: ECB’s Area-Wide Model database.

(i.e. total tax revenues minus transfers), the wage rate, all in real per capita terms, and the annual rate of change of the Harmonized Index of Consumer Prices (HICP).\textsuperscript{12} Those additional variables are added all together in the extended specification of the fixed parameters model. In the specification of the model with time-varying parameters we are however constrained by the need to avoid overparameterization and exhausting available degrees of freedom. The additional variables are therefore added one at a time to the benchmark specification, thus limiting the number of variables in the TVP-VAR

\textsuperscript{12}We use the HICP-based inflation rate to assess the response of inflation to spending shocks due to its close link to monetary policy decisions in the euro area.
to a maximum of five indicators.

As Burriel et al. (2010), we use a quarterly fiscal data set for the euro area compiled by Paredes et al. (2009). The latter employ mixed-frequencies state space models estimated with available (mostly annual) national accounts data and monthly and quarterly fiscal information taken from government cash accounts to obtain interpolated quarterly fiscal data for the above-mentioned period. By construction, the interpolated variables are coherent with official ESA95 annual and quarterly euro area data, as far as the latter is available. This approach has the advantage that it avoids the endogenous bias which could arise if fiscal data interpolated on the basis of general macroeconomic indicators were used with macroeconomic variables to assess the impact of fiscal policies. Other macroeconomic data for the euro area are mainly taken from the ECB’s Area-Wide Model database (see Fagan et al., 2005).

To ensure comparability with the previous literature, our data definitions closely follow related studies. Details are provided in Appendix 2.D. Figure 2.1 shows the data used in the benchmark specification. Both models are estimated in levels and prior to the estimation all variables except the interest rate and the inflation rate were transformed into natural logarithms.

### 2.3.4 Structural interpretation

The reduced-form models attempt to capture structural representations with uncorrelated shocks. The reduced-form innovations are therefore linear transformations of some underlying structural shocks $e_t$ with $E[e_t e'_t] = I$, i.e. $u_t = C e_t$ for the fixed parameters model and $u_t = C_t e_t$ for the model with time-varying parameters, for $t = 1, 2, 3, \ldots, T$. In particular, the innovations in the equation for government spending can be considered as linear combinations of three types of shocks (see Blanchard and Perotti, 2002): (i) the automatic response of spending to movements in the business cycle, prices and interest rates; (ii) the systematic discretionary response of spending to macroeconomic developments; (iii) deliberate discretionary changes in spending. The latter are the truly structural spending shocks of interest.

Without restrictions on the matrices $C$ and $C_t$, and therefore the covariance matrices $R$ and $R_t$, the above systems are not identified since many combinations of struc-
tural shocks can generate the same reduced-form innovations. To achieve a structural representation, government spending shocks are identified by assuming that government spending is predetermined in a system with output, consumption, the interest rate, and possibly other macroeconomic variables, following Blanchard and Perotti (2002) and Fatás and Mihov (2001). Thus, government spending is ordered first in the estimated models and the desired linear combination is achieved by a Cholesky decomposition, i.e. $R = CC'$ and $R_t = C_tC_t'$, where $C$ and $C_t$ are lower triangular matrices. Under this recursive identification scheme, all variables are allowed to respond within a quarter to innovations to government spending but government spending does not react within a quarter to innovations to other variables in the system.

As discussed by Caldara and Kamps (2008), the fact that government spending as defined here does not include interest payments justifies that spending is ordered before the interest rate. The fact that spending is defined net of transfer payments further justifies the assumption of acyclical, i.e. there is no automatic contemporaneous reaction of spending to movements in the business cycle. In addition, due to implementation lags in the policy process, an immediate discretionary fiscal response to a change in the economy is unlikely to occur. When more variables are included, the assumption that government spending does not react within a quarter to shocks to those variables can be justified on similar grounds.

As mentioned above, a well-known criticism of the above SVAR approach centers on the possibility that autonomous policy changes can be anticipated by economic agents (see e.g. Ramey, 2011b; Leeper, Walker, and Yang, 2011). This criticism is addressed in Section 2.5, based on Granger-causality tests that relate the identified SVAR shocks to institutional forecasts and survey data, following Ramey (2011b).

Impulse responses are then calculated as follows. In the fixed parameters case, given draws from the posterior distributions of $R = CC'$ and the $B_i$, the first column of the matrix $C$ gives the contemporaneous responses (at horizon $k = 0$) of the endogenous variables to a one-time, unitary structural shock to government spending $e_0 = [1, 0, \ldots, 0]'$, and model (2.1) with $u_k = [0, 0, \ldots, 0]'$ can be used to calculate impulse responses at horizons $k \geq 1$. In the time-varying parameters case, we apply a local approximation to the impulse responses at time $t$, following e.g. Gali and Gam-
betti (2009). That is, the contemporaneous responses to unitary shocks $e_{t,0}$ at time $t$ are derived from draws from the posterior distribution of reduced-form covariance matrices $R_t = C_t C_t'$, and the draws from the distribution of the $B_{i,t}$ are applied to calculate impulse responses at horizons $k \geq 1$, using model (2.2) with $u_{t,k} = 0$.

2.4 Estimated effects of spending shocks

In this section, we first present estimation results for the identified fixed parameters model, to assess the impact of government spending shocks over the full sample. We then discuss results for the identified time-varying parameters model.

2.4.1 Time-invariant impulse responses

Figure 2.2 reports the estimated impulse responses due to the identified government spending shocks to the four endogenous variables $y_t$ of equation (2.1) for the benchmark specification, together with their 16 and 84 percent probability bands. The responses of output, consumption, and spending (and later on investment and net taxes) to the spending shock are reported as non-accumulated multipliers. That is, the original impulse responses are divided by the impact response of government spending and the result is divided by the ratio of government spending and the responding variable. The rescaled impulse responses can thus be interpreted to give the reaction of the responding variable, in percent of real GDP, to a spending shock leading to an initial increase in the level of government spending of size 1% of real GDP. For the fixed parameters model, the ratio is evaluated at the sample mean. For the model with time-varying parameters, we take the ratio in the respective quarter.\footnote{For example, suppose that the shock leads to a two percent increase in government spending. Since the share of spending over GDP is roughly 25 percent, this corresponds to a spending increase of about 0.5% of GDP. Say output increases by one percent and consumption increases by 0.5 percent, i.e. by 0.25% of GDP since the share of consumption over GDP is about 50 percent. The share of spending over consumption is thus roughly 50 percent. The multipliers would be $(1/2)/0.25 = 2$ for output and $(0.5/2)/0.5 = 0.5$ for consumption.}

According to the results in Figure 2.2, a government spending shock induces a positive, persistent response of spending lasting more than four years. The initial reaction of output is positive, the estimated short-run multiplier being 0.54. The
Notes. Median impulse responses with 16 and 84 percent probability bands; the responses of output, consumption, and spending are measured in percent of GDP to a 1% of GDP spending shock, i.e transformed response at horizon $k = \text{responding variable’s original response at horizon } k/(\text{spending response at horizon } 0 \times \text{average ratio of spending to responding variable});$ the response of the interest rate is measured in percentage points to a one percent shock.

output response remains positive with 68 percent probability for about one year after the shock, it turns negative after two years, it reaches a minimum after about three years, and it then returns to the baseline. The spending shock also leads to a positive initial response of private consumption. Similarly as for output, however, the response of consumption turns negative over the medium term. The nominal interest rate hardly responds to the spending shock on impact, but it then starts to rise and peaks about one year after the shock. The interest rate response is estimated to be positive with 68 percent probability from two quarters until around three years after the shock.

In a next step we extend the VAR specification by a broader set of indicators
Figure 2.3: Time-invariant impulse responses II – extended specification

Notes. See Figure 2.2; the responses of output, consumption, investment, spending, and net taxes are measured in percent of GDP to a 1% of GDP spending shock; the response of the real wage is measures in percent to a one percent shock; the responses of the interest rate and inflation are measured in percentage points to a one percent shock.

which often appear in related studies. The estimated impulse responses of government spending, output, consumption, investment, the real wage, net taxes, the HICP-based inflation rate, and the nominal interest rate are reported in Figure 2.3. As a consequence of a 1% of GDP spending increase, net taxes increase by about 0.8% of GDP on impact, indicating an overall fiscal expansion since the primary deficit increases. Net taxes also return more quickly to baseline than spending does, thus the shock remains expansionary. Output, consumption, and investment increase at first but fall afterwards below their initial levels. The responses of output and the components of domestic private demand are however estimated with relatively little precision. The
point estimates of the impact multipliers are 0.55 (output), 0.23 (consumption) and 0.03 (investment).\footnote{An output multiplier smaller than one combined with (marginally) positive point estimates for consumption and investment could be explained by a decline in net exports, although we have not included this variable as it is not available at the euro area level. This explanation is however consistent with SVAR results for a panel of 14 EU countries discussed in Beetsma, Giuliodori, and Klaassen (2008), showing that, on average, the trade balance falls by 0.5% of GDP on impact due to a 1% of GDP increase in government spending.} The real wage increases by approximately 0.15 percent on impact and remains above its initial level during more than three years after the shock. Inflation shows a muted response in the initial two quarters but it starts to increase later on. The nominal interest rate reacts similarly as in the benchmark specification.

Overall, these results indicate that on average over the period 1980-2008 government spending shocks have had expansionary short-run effects on output, consumption, investment, and real wages in the euro area. However, output declines at longer horizons as consumption and investment are being crowded out. The estimated increase in the nominal interest rate is consistent with an offsetting reaction of monetary policy to the fiscal expansion to reduce inflationary pressures. At the same time, our findings compare well with the results of previous SVAR studies for the euro area. In particular, they are broadly similar to the results of Burriel et al. (2010), the main previous fiscal VAR study for the euro area employing a similar data set. Burriel et al. (2010) also find a positive impact of government spending shocks on GDP and private consumption in the short run and a decline at longer horizons, an increase in the aggregate primary government deficit, and a relatively persistent increase in interest rates.

\subsection{Time-varying impulse responses}

The time-varying nature of model (2.2) allows to compute state-dependent impulse responses for individual quarters of the estimation sample. In the following, we look at the results from various different perspectives.

Figure 2.4 shows the estimated impulse response functions for the variables in the benchmark specification for three selected quarters at the beginning (1980Q4), towards the middle (1995Q4), and at the end of the sample (2008Q4). The results show that the estimated short-run multiplier on output is larger at the beginning of the sample, the point estimate being around 0.7 for 1980Q4 compared to 0.4 for 2008Q4. Moreover,
Figure 2.4: Time-varying impulse responses I – selected quarters

Notes. Median impulse responses with 16 and 84 percent probability bands; the responses of output, consumption, and spending are measured in percent of GDP to 1% of GDP spending shocks, i.e. transformed response at horizon \( t \) and horizon \( k = \) responding variable’s original response at time \( t \) and horizon \( k = 0 \times \) ratio of spending to responding variable at time \( t \); the responses of the interest rate are measured in percentage points to one percent spending shocks.

The effects of spending shocks on output seem to have lost persistence over time, and they are increasingly negative at longer horizons. In particular, the estimated response of GDP at a horizon of five years is about -0.7 percent for 1980Q4 but -1.6 percent for 2008Q4. The time-varying techniques thus indicate increasingly contractionary longer-term effects of a spending expansion. Furthermore, while the initial output response is positive with 68 percent probability in the initial period, the probability bands include the zero line at the end of the sample. Instead, the response after five years is significantly negative only in the most recent period. The results further suggest that the effects on consumption have decreased over time in a similar way as the effects on
output. We also note a stronger response of the nominal interest rate.

The conclusions from Figure 2.4 are confirmed in Figure 2.5, which shows the estimated state-dependent median impulse responses for each year in the sample. Only the fourth-quarter response in each year is reported, such that the first impulse response refers to 1980Q4 while the last one refers to 2008Q4. The figure shows that the estimated short-run effects of spending shocks on output and consumption are largest towards the end of the 1980s and lowest towards the recent period, whereas the estimated effects at longer horizons are steadily falling from the beginning towards the end of the sample. The estimated impulse response of government spending itself is however rather stable over time.
Figure 2.6: Time-varying impulse responses III – selected horizons

Notes. See Figure 2.4; \( k \)-th horizon median responses for \( k = 0, 4, 20 \).

Figure 2.6 shows the responses of all variables over time at selected horizons, i.e. the contemporaneous responses, the responses after one year, and the responses after five years. The estimated contemporaneous multiplier on output is slightly below one for the period 1980-1985, larger than one for the period 1986-1990, and then falls over the period 1991-2003 to reach values around 0.5 in 2004-2008. At the five-year horizon, the estimated effects on output and consumption of an initial 1% of GDP expansion are substantially lower for the recent decade, from -1.4% to -1.7% of GDP, compared to -0.7% to -1% in the 1980s. In general, the changes in the effects on output are similar as the changes in the effects on private consumption. The estimated contemporaneous reaction of the nominal interest rate is negative from 1980 until around 1999-2002.
Figure 2.7: Pair-wise joint posterior distributions of time-varying impulse responses

Notes. See Figure 2.4; pair-wise responses of output, consumption (in percent of GDP), and the interest rate (in percentage points), computed across their posterior distribution at horizons $k = 0, 4, 20$. and positive afterwards. The estimated interest rate response at longer horizons also increases over time. A stronger interest rate response thus might have contributed to the decrease in fiscal multipliers.

To test differences in the above responses over time, we compute the joint pair-wise distributions of impulse responses at two selected horizons. That is, in Figure 2.7 (sorted) draws from the posterior distribution of output and consumption responses and the interest rate response in 1980Q4 are plotted against draws for 2008Q4.\textsuperscript{15} Results are

\textsuperscript{15}A similar exercise is implemented in Cogley et al. (2010). There are many alternative pairs of quarters to choose from, but the results are not particularly sensitive to this choice as long as the
Figure 2.8: Time-varying impulse responses IV – extended specifications

Notes. See Figure 2.4; only median impulse responses are reported.

reported for the impact responses, the one-year responses, and the five-year responses. Each point in the respective panels represents a draw from the joint distribution for 1980Q4 and 2008Q4. Thus, combinations near the 45 degree line represent pairs for which there was little or no change over time and those above (below) the 45 degree line are pairs where the response of the respective variable has increased (decreased). The figure shows that the lower tails of the distributions of the output and consumption responses have shifted downwards, especially at longer horizons, whereas the upper tails appear comparably stable. Therefore, the median estimates have shifted downwards as well. Regarding the interest rate, both time variation in its impact response and the response after five years turn out to be important.

periods used are sufficiently distant from each other.
We also investigate time variation in the effects of spending shocks on a broader set of macroeconomic indicators, adding one at a time private investment, net taxes, the real wage, and the HICP-based inflation rate to the estimated VAR. Figure 2.8 shows the estimated state-dependent median impulse responses. We observe a small positive short-term effect on private investment and a crowding out at longer horizons. Similarly as the multiplier on output and the effect on consumption, the effect on investment was larger in the first part of the sample. The reaction of net taxes to government spending shocks has remained comparably stable over time, and throughout the response is smaller than 1% of GDP, indicating that the primary deficit has always increased due to the spending shock. A smaller overall fiscal expansion can thus not serve as an explanation for smaller spending multipliers.

The response of the real wage is estimated to be positive for several quarters after the shock throughout the sample, but it shows a larger initial reaction and a more persistent response in the first part of the sample towards the late 1980s. The initial response of inflation was close to zero throughout, but we observe a stronger medium-term response during the 1980s and most of the 1990s. As the nominal interest rate reacts more strongly to government spending shocks, this result implies that the real interest rate response has tended to increase over time.

2.5 Robustness checks

This section reports the results of several robustness checks, as listed below.

2.5.1 Scaling factors

In the estimation of the TVP-VAR, we have elicited relative conservative priors on time variation, in particular the scaling factors $k_Q$, $k_S$, and $k_W$ which parameterize the priors on the covariance matrices of the shocks in the state equations, as described in Appendix 2.B. The values were $k_Q = k_W = 10^{-4}$ and $k_S = 10^{-2}$, following the related literature (see, in particular, Primiceri, 2005). To check the sensitivity of the estimation outcomes, we now further reduce the scaling factors one at a time to $k_Q = k_W = 0.5 \times 10^{-4}$, and $k_S = 0.5 \times 10^{-2}$, keeping the other two factors fixed at
Figure 2.9: Robustness I – smaller prior scaling factor $k_Q$

![Graphs showing changes in Government Spending, Output, Consumption, and Interest Rate over time.](image)

Note. See Figure 2.6.

The results are summarized in Figures 2.9, 2.10, and 2.11, respectively. Figure 2.9 shows that the reduction of the coefficients scaling factor $k_Q$ especially increases the estimated time variation of the short-run multiplier on output and the contemporaneous effect on consumption. It therefore seems that, compared to the previous results, some of the time variation in the VAR coefficients is instead picked up by the covariance terms. On the other hand, Figures 2.10 and 2.11 show that the reductions of $k_S$ and $k_W$ only lead to relatively small changes in the amount and the direction of the estimated time variation.
Figure 2.10: Robustness II – smaller prior scaling factor $k_S$

Note. See Figure 2.6.

2.5.2 Stationarity conditions

In their analysis of U.S. monetary policy, Cogley and Sargent (2002) have proposed to discard draws from the Gibbs sampler that do not satisfy stationarity conditions, and many related studies have followed this approach. However, we have argued above that the stationarity restriction is harder to defend for aggregate euro area fiscal data since fiscal variables may have followed unstable paths in some countries. The potential downside of not imposing the stationarity conditions is that this may exaggerate the amount of time variation due to a potentially large amount of unstable draws. We therefore check the robustness of the TVP-VAR results when stationarity conditions are
imposed on the VAR coefficients. Formally, the random walk process 2.3 for the VAR coefficients $\beta_t$, $t = 1, 2, 3, \ldots, T$, characterizes the conditional density $f(\beta_t|\beta_{t-1}, Q)$. Following Cogley and Sargent (2002), introduce an indicator function $I(\beta_t)$ which rejects unstable draws that do not satisfy standard eigenvalue stability conditions and which thus enforces stationarity of the estimated TVP-VAR at each point of time. The VAR coefficients are thus postulated to evolve according to

$$p(\beta_t|\beta_{t-1}, Q) = I(\beta_t)f(\beta_t|\beta_{t-1}, Q).$$

Figure 2.12 shows the estimated state-dependent effects at selected horizons when
the stationarity conditions are imposed. A comparison with the previous results indicates no significant differences to the benchmark case. The multipliers show somewhat less high-frequency variation but the broad patterns are similar.

### 2.5.3 Anticipation effects

To check for the possible presence of anticipation effects, this section confronts the estimated SVAR shocks with macroeconomic forecasts to see whether the identified shocks are potentially predictable. This exercise follows Ramey (2011b) who shows that, for the U.S., SVAR spending shocks are Granger-caused by forecasts made one
to four quarters earlier (i.e. they are predictable). Thus, we perform tests of Granger causality from various variables conveying information about future policy and macro developments onto the time series of estimated SVAR spending shocks. We use both survey data from Consensus Economics, as in Ramey (2011b), and publicly available short-term forecasts by the European Commission. The Consensus data summarizes the predictions of professional forecasters at banks and other financial institutions. This data is thus taken to represent economic agents’ (or market participants’) expectations on future macroeconomic developments. The European Commission forecasts do not directly reflect such expectations, but they do cover a longer period than the survey data, thus increasing the power of the tests.\footnote{The European Commission provides forecasts in November of every year for the following year since the 1970s for a number of European countries. Consensus Economics provides forecasts every month for 1991 onwards. Forecasts on the budget deficit are only available for 1994 onwards.} We therefore exploit both data sets.

The exercise conducted below is however subject to the following limitations. First, we are forced to use time-aggregated quarterly data in the estimation since the macroeconomic forecasts are only available on an annual basis. Second, we also need to restrict the analysis to the fixed parameters VAR as the number of observations in the time-aggregated data is not sufficient to estimate the TVP-VAR. Third, the data incorporates predictions on government deficits and deficit-to-GDP ratios as the only fiscal variables instead of direct forecasts on government spending. The results reported below should be interpreted with these limitations in mind.

The results of the Granger tests are reported in Table 2.1. Following Ramey (2011b), the SVAR shocks in period $t$ are regressed on a constant, their own lags and various forecasts made in period $t-1$ for period $t$.\footnote{The results are robust to the use of additional lagged values of the left-hand side and/or the right-hand side variables, as well as the addition of the period $t$ variables (and their lagged values) which are included in the VAR model on the right-hand side.} The null hypothesis is that the forecasts do not Granger-cause the SVAR shocks.\footnote{The Granger causality test is identical to an F-test of the null hypothesis that the unrestricted model, which includes the forecasts, does not provide a better fit than the restricted model, which excludes the forecasts.} The first panel of Table 2.1 shows that the null hypothesis cannot be rejected at the 10 percent significance level for any of the European Commission’s forecasts in isolation, on the deficit-to-GDP ratio and real GDP growth, and also not if both forecasts are included as right-hand side variables. Similarly, the second panel shows that the null hypothesis cannot be
Table 2.1: Granger causality tests using macroeconomic forecasts\(^a\)

<table>
<thead>
<tr>
<th>Hypothesis test</th>
<th>F-statistic</th>
<th>10% critical value</th>
<th>Conclusion (p-value)</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>European Commission Forecasts</strong> c</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Deficit-to-GDP ratio forecasts → SVAR shocks</td>
<td>0.004</td>
<td>2.949</td>
<td>No (0.949)</td>
</tr>
<tr>
<td>GDP growth forecasts → SVAR shocks</td>
<td>0.001</td>
<td>2.949</td>
<td>No (0.973)</td>
</tr>
<tr>
<td>All forecasts → SVAR shocks</td>
<td>0.002</td>
<td>2.575</td>
<td>No (0.998)</td>
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<tr>
<td>Deficit-to-GDP ratio forecasts → actual spending growth</td>
<td>4.894</td>
<td>2.949</td>
<td>Yes (0.038)</td>
</tr>
<tr>
<td><strong>Consensus Economics Forecasts</strong> c</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Deficit growth forecasts → SVAR shocks</td>
<td>0.027</td>
<td>3.225</td>
<td>No (0.872)</td>
</tr>
<tr>
<td>GDP growth forecasts → SVAR shocks</td>
<td>0.373</td>
<td>3.102</td>
<td>No (0.551)</td>
</tr>
<tr>
<td>Consumption growth forecasts → SVAR shocks</td>
<td>0.155</td>
<td>3.102</td>
<td>No (0.700)</td>
</tr>
<tr>
<td>Interest rate forecasts → SVAR shocks</td>
<td>0.785</td>
<td>3.102</td>
<td>No (0.391)</td>
</tr>
<tr>
<td>All forecasts → SVAR shocks</td>
<td>0.049</td>
<td>2.693</td>
<td>No (0.995)</td>
</tr>
<tr>
<td>Deficit growth forecasts → actual spending growth</td>
<td>0.320</td>
<td>3.225</td>
<td>No (0.320)</td>
</tr>
</tbody>
</table>

\(^a\) The first variable at time \(t\) is regressed on a constant, its own lag at time \(t - 1\), and the forecast made at time \(t - 1\) of the second variable for period \(t\).

\(^b\) The null hypothesis is that the second variable does not Granger-cause the first variable.

\(^c\) For the European Commission forecasts (1982-2006), GDP is measured as real annual growth rate and the deficit-to-GDP ratio is measured in nominal terms. For the Consensus Economics forecasts (1992-2008), all variables except the interest rate are measured as real annual growth rates, using consumer prices as deflators, and the interest rate is measured in nominal terms. See Appendix 2.D for details on the data definitions.
rejected for the professional forecasts on the growth rates of the budget deficit, GDP, private consumption, and the short-term interest rate. In addition, we check whether the Commission’s forecasts on the deficit-to-GDP ratio and professional forecasts on the budget deficit Granger-cause realized spending growth. This is the case for the Commission’s forecasts, where the null hypothesis is rejected. Thus, although the forecasts do have predictive power for realized spending, they do not predict the SVAR spending shocks. Overall, this exercise does not provide strong reasons to doubt the validity of the identification approach due to anticipation effects.

2.6 The fiscal transmission mechanism

This section exploits the results obtained in the previous step with the aim of providing empirical evidence on the determinants of the effects of government spending shocks in the euro area. Section 2.6.1 reviews the main theories on the fiscal transmission mechanism, focusing on (i) the level of government debt, (ii) asset market participation and the availability of credit, (iii) the degree of trade openness, (iv) the share of government investment in total spending, and (v) the wage component of total spending. Section 2.6.2 relates these factors to the estimated effects of spending shocks using regression analysis.

2.6.1 Views on the transmission mechanism

(i) Government debt. Experience from past fiscal consolidations suggests the possibility that in times of fiscal stress an economy’s response to fiscal shocks changes. That is, positive consumption growth was observed after prolonged and substantial deficit cuts. This is the hypothesis of “expansionary fiscal contractions” brought about by Giavazzi and Pagano (1990).\(^\text{19}\) Indeed, for a panel of 19 OECD countries, Perotti (1999) finds that the effect of spending shocks on consumption can be positive if the

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\(^{19}\)See also Giavazzi, Jappelli, and Pagano (2000). Giavazzi and Pagano (1990) study episodes of large fiscal consolidations in Denmark during 1983-1986 and in Ireland during 1987-1989. In these episodes the cyclically adjusted deficit as a share of GDP declined by 9.5 percent and 7.2 percent relative to the preconsolidation year and yet private consumption increased by 17.7 percent and 14.5 percent cumulatively. Alesina and Perotti (1996) identify similar episodes in several other European countries and Canada during the 1980s.
initial financing needs of the government are small, arguing that this outcome is due to the convexity of tax distortions: a (larger) expected increase in taxation tomorrow causes a (larger) decline in wealth and a (larger) fall in consumption today.

(ii) Credit. In standard general equilibrium models, expansionary government spending shocks tend to generate a crowding out of private consumption. The reason is the negative wealth effect induced by higher future tax payments, which increases consumer saving due to the consumption smoothing objective. However, credit constraints and limited asset market participation may dampen this effect. For example, Galí et al. (2007) show that a spending shock can generate an increase in aggregate consumption in a New Keynesian model conditional on, in particular, a relatively large fraction of liquidity-constrained consumers. In addition, it has recently been argued that fiscal stabilization policy may be more effective during recessions since credit constraints might then bind across a wider range of agents. In particular, Roeger and in ’t Veld (2009) allow for credit-constrained households along the lines of the financial accelerator literature, thus allowing the stringency of credit constraints to vary over the cycle, and show that stabilization policy becomes more effective since the propensity to consume out of current income increases during recessions.\(^{20}\)

(iii) Openness. It is often claimed that the effectiveness of fiscal policy depends on the degree of openness to trade.\(^{21}\) The argument is that in very open economies domestic output will be comparatively less affected by a fiscal expansion since a large fraction of the intended stimulus falls on imports. For instance, Beetsma, Giuliodori, and Klaassen (2008) show that a 1% of GDP increase in public spending in the EU leads to a fall of the trade balance by 0.5% of GDP on impact and a peak fall of 0.8% of GDP. With respect to time variation in fiscal multipliers, the effects of an increase in spending on GDP are then expected to be smaller the higher the degree of openness. Below we use the import share as a proxy for the degree of openness since imports

\(^{20}\)Tagkalakis (2008) also provides evidence for asymmetric effects of fiscal policy for a panel of 19 OECD countries over the period 1970-2002, showing that a spending shock has a larger effect on private consumption in downturns than in upturns. See also Auerbach and Gorodnichenko (2010).

\(^{21}\)See, for instance, Perotti (2005) who however argues that the increase in openness is probably too small to account for the decline in spending multipliers in OECD economies.
are the relevant channel through which openness to trade may affect fiscal multipliers according to this argument.

(iv) Government investment. Although not all empirical studies find a growth-enhancing effect of public capital, there is now more consensus than in the past that public capital supports economic growth (see Romp and de Haan, 2007). A corresponding change in the composition of spending may therefore contribute to changing spending multipliers. Macroeconomic models which account for productive public capital typically predict that increases in government investment can generate larger fiscal multipliers than increases in government consumption, due to the beneficial aggregate supply effect of public capital.\footnote{See, for instance, Baxter and King (1993), Pappa (2005), and Straub and Tchakarov (2007).} On the other hand, Leeper, Walker, and Yang (2010) have recently provided evidence showing that government investment projects in the U.S. are subject to substantial implementation lags. Private investment and employment are then postponed until the public capital is on line, which, as these authors show in a macroeconomic model, can lead to smaller short-term multipliers.

(v) Wage component. More than half of government consumption in the euro area consists of wage payments to government employees, whereas less than half consists of goods purchases. Several studies emphasize that this distinction matters when assessing the impact of spending shocks on the macroeconomy. For example, based on a neoclassical model, Finn (1998) shows that government employment shocks raise the real wage and thus act as a transfer to households, which dampens the wealth effects on consumption and labor supply.\footnote{Pappa (2005) demonstrates that government employment shocks have similar effects in a New Keynesian framework.} Using SVAR analysis, Perotti (2008) shows that, in the U.S., government employment shocks have larger effects on output and consumption than shocks to government goods purchases. On the other hand, Alesina and Ardagna (2010) argue that in an imperfect labor market a decrease in government employment could reduce job finding probabilities, whereas a decrease in government wages could decrease incomes of workers in the public sector. In both cases, reservation utilities and wages demanded for private sector workers would decrease, which
Chapter 2

may increase profits, investment, and competitiveness.

2.6.2 Driving forces of time variation

Several testable hypotheses can be derived from the discussion in Section 2.6.1. First, the effects of spending shocks on output and consumption are expected to be smaller the higher the initial debt-to-GDP ratio. Second, the effects can be higher if households are more restricted in their access to credit, or if actual output is below potential output. Third, a higher share of imports over GDP is expected to lead to smaller spending multipliers. Fourth, a higher government investment share can lead to higher spending multipliers, but if implementation lags play a role, short-term multipliers can also be smaller. Fifth, a higher wage share can result in larger or smaller effects on economic activity depending on the degree of labor market competitiveness.

The above hypotheses are analyzed using regression inference. We apply Bayesian linear regressions, using the estimated time-varying effects on output and consumption as dependent variables. This type of two-step approach, while based here on time-varying parameters, is close in spirit to Fatás and Mihov (2006). We distinguish both contemporaneous effects and longer-term effects after five years. Further, since the dependent variables are themselves estimated parameters, the standard errors of the regression coefficients are adjusted to account for the uncertainty in the dependent variables. Not doing so may give a biased view on the importance of the restrictions implied by the explanatory variables and might thus artificially produce significant effects even when the “true” ones are negligible (see Canova and Pappa, 2006). In particular, we use each of 1,000 posterior draws of multipliers from the identified TVP-VAR model in turn as dependent variable. We then generate 1,100 draws of regression coefficients by Gibbs sampling and omit the first 100 draws for each regression. This

\footnote{Diffuse normal priors with mean zero and standard deviation 10^6 are specified for the regression coefficients. All regressions include a constant and a linear trend to address possible concerns of spurious causation. Using a linear-quadratic trend instead of a linear trend did not lead to significant changes in the results. We also account for the possible presence of heteroskedastic disturbances, where we use diffuse priors on the variance terms. The regressions are estimated using a Gibbs sampling algorithm with 1,100 draws, dropping the first 100 draws, see Geweke (1993).}

\footnote{Fatás and Mihov (2006) study the determinants of output elasticities of government spending. The latter are estimated in a first step over a sample of 48 U.S. states. In a second step, the authors analyze the impact of different fiscal rules on those elasticities.}
Figure 2.13: Potential determinants of spending multipliers

Notes. The debt-to-GDP ratio is in nominal annual terms; the ratio of credit to households over GDP is outstanding (end-of-period) loans to households divided by the sum of nominal GDP of the last four consecutive quarters; the output gap is measured as quarterly percentage deviation from trend real GDP, trend is based on HP filter with smoothing parameter 1600; the ratio of imports over GDP and the shares of government investment and wage expenditures in total spending are based on quarterly nominal data; source of fiscal data: Paredes et al. (2009); source of remaining data: ECB’s Area-Wide Model database and Bank of International Settlements macroeconomic series (data on loans).

leaves us with 1,000,000 posterior draws from the posterior distribution of regression coefficients, conditional on the full posterior distribution of estimated multipliers, from which we compute means and posterior probabilities.

Figure 2.13 shows the explanatory variables used in the regression analysis. The lagged aggregate euro area debt-to-GDP ratio is used to measure the initial financing needs of euro area governments. Availability of credit is measured by the lagged ratio of credit to households over GDP. The state of the business cycle is approximated by the lagged HP-filtered output gap. Lagged values are used to address potential reverse causation from spending multipliers on output and the business cycle. The ratio of
imports over GDP (in lagged terms) is used to assess the impact of changes in the
degree of openness. Finally, we include the contemporaneous shares of government
investment and employee compensation over total spending to assess the impact of
changes in the composition of spending on its overall effects.

The results for contemporaneous effects and the effects after five years, respectively,
are reported in Tables 2.2 and 2.3. The point estimates of the regression coefficients are
the means of their posterior distribution. The statistical significance of the regression
coefficients is measured as the posterior probability that they are non-positive (non-
negative) if their point estimates are positive (negative).

The results in Table 2.2 show that, on average, an increase in the debt-to-GDP ratio
has a negative but small effect on short-term multipliers. On the other hand, a rise in
the credit ratio is estimated to have a larger impact, a one percentage point increase
leading on average to a decline in the spending effect on output (consumption) between
0.04 and 0.06 points (between 0.02 and 0.04 points). The credit ratio has increased
from 30 percent in 1980 to almost 60 percent in 2008, such that increasing credit avail-
ability is estimated to have contributed substantially to the decline in contemporaneous
multipliers. The output gap enters with an unexpected positive sign, whereas a rise
in the import share is estimated to have a negative but mostly insignificant effect.
The estimated impact of an increase in the share of government investment in total
spending is positive whereas an increase in the share of wage payments is estimated
to have a negative effect. In the largest regression model for output (consumption), a
unitary increase of the investment share is estimated to cause an average increase in
the contemporaneous effects by 0.07 points (0.04 points). A unitary increase in the
wage share leads to a decrease in the effects by 0.04 points (0.03 points).

The evidence presented in Table 2.3, on the other hand, suggests that the level of
government debt relative to GDP is the main determinant of the longer-term effects
of government spending. For both output and consumption, a one percentage point
increase in the debt ratio leads on average to a decline by 0.01 points in the associated
effects, the coefficients being negative with at least 95 percent probability in all regres-
sion models. The coefficients on some of the remaining variables do have the expected
signs, but none of them are different from zero with more than 90 percent probability.
Table 2.2: Bayesian linear regressions on contemporaneous effects\textsuperscript{a,b}

<table>
<thead>
<tr>
<th>Multiplier on output</th>
<th>Effect on consumption</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
</tr>
<tr>
<td>Gov. debt/GDP (-1)</td>
<td>0.01</td>
</tr>
<tr>
<td></td>
<td>(0.01)</td>
</tr>
<tr>
<td>Credit/GDP (-1)</td>
<td>-0.06***</td>
</tr>
<tr>
<td></td>
<td>(0.02)</td>
</tr>
<tr>
<td>Output gap (-1)</td>
<td>0.03**</td>
</tr>
<tr>
<td></td>
<td>(0.02)</td>
</tr>
<tr>
<td>Imports/GDP (-1)</td>
<td>-0.02*</td>
</tr>
<tr>
<td></td>
<td>(0.01)</td>
</tr>
<tr>
<td>Inv. share</td>
<td>0.03*</td>
</tr>
<tr>
<td></td>
<td>(0.02)</td>
</tr>
<tr>
<td>Wage share</td>
<td>-0.04**</td>
</tr>
<tr>
<td></td>
<td>(0.03)</td>
</tr>
<tr>
<td>Constant</td>
<td>0.71</td>
</tr>
<tr>
<td></td>
<td>(1.37)</td>
</tr>
<tr>
<td>Trend</td>
<td>-0.01***</td>
</tr>
<tr>
<td></td>
<td>(0.00)</td>
</tr>
<tr>
<td>Observations</td>
<td>112</td>
</tr>
</tbody>
</table>

\textsuperscript{a} The Bayesian regressions allow for heteroskedastic errors following Geweke (1993). The standard error adjustment proceeds by using each of 1,000 multipliers in the posterior distribution from the identified TVP-VAR as dependent variable. All regressions are then estimated using a Gibbs sampling algorithm with 1,100 draws and 100 omitted draws. This leaves us with 1,000,000 posterior draws of regression coefficients.

\textsuperscript{b} The point estimates are the posterior means of the posterior distribution. Standard deviations are reported in parentheses. Asterisks indicate posterior probabilities that the regression coefficients are non-positive if the point estimates are positive or non-negative if the point estimates are negative (*less than ten percent, **less than five percent, ***less than one percent). Explanatory variables are measured in percent.
### Table 2.3: Bayesian linear regressions on effects after five years\textsuperscript{a,b}

<table>
<thead>
<tr>
<th></th>
<th>Multiplier on output</th>
<th>Effect on consumption</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
</tr>
<tr>
<td>Gov. debt/GDP (-1)</td>
<td>-0.00</td>
<td>-0.01**</td>
</tr>
<tr>
<td></td>
<td>(0.02)</td>
<td>(0.02)</td>
</tr>
<tr>
<td>Credit/GDP (-1)</td>
<td>-0.02</td>
<td>-0.02</td>
</tr>
<tr>
<td></td>
<td>(0.09)</td>
<td>(0.09)</td>
</tr>
<tr>
<td>Output gap (-1)</td>
<td>-0.01</td>
<td>-0.02</td>
</tr>
<tr>
<td></td>
<td>(0.10)</td>
<td>(0.11)</td>
</tr>
<tr>
<td>Imports/GDP (-1)</td>
<td>0.01</td>
<td>0.01</td>
</tr>
<tr>
<td></td>
<td>(0.03)</td>
<td>(0.04)</td>
</tr>
<tr>
<td>Inv. share</td>
<td>-0.00</td>
<td>-0.01</td>
</tr>
<tr>
<td></td>
<td>(0.13)</td>
<td>(0.12)</td>
</tr>
<tr>
<td>Wage share</td>
<td>0.01</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(0.08)</td>
<td></td>
</tr>
<tr>
<td>Constant</td>
<td>-0.67</td>
<td>0.02</td>
</tr>
<tr>
<td></td>
<td>(1.96)</td>
<td>(3.24)</td>
</tr>
<tr>
<td>Trend</td>
<td>-0.01***</td>
<td>-0.01***</td>
</tr>
<tr>
<td></td>
<td>(0.02)</td>
<td>(0.01)</td>
</tr>
<tr>
<td>Observations</td>
<td>112</td>
<td>112</td>
</tr>
</tbody>
</table>

\textsuperscript{a} The Bayesian regressions allow for heteroskedastic errors following Geweke (1993). The standard error adjustment proceeds by using each of 1,000 multipliers in the posterior distribution from the identified TVP-VAR as dependent variable. All regressions are then estimated using a Gibbs sampling algorithm with 1,100 draws and 100 omitted draws. This leaves us with 1,000,000 posterior draws of regression coefficients.

\textsuperscript{b} The point estimates are the posterior means of the posterior distribution. Standard deviations are reported in parentheses. Asterisks indicate posterior probabilities that the regression coefficients are non-positive if the point estimates are positive or non-negative if the point estimates are negative (*less than ten percent, **less than five percent, ***less than one percent). Explanatory variables are measured in percent.
To summarize, the second-stage regressions indicate that (i) a higher level of government debt relative to GDP is associated with lower spending multipliers at longer horizons. (ii) The ratio of credit over GDP seems to be an important determinant of the observed time variation in the short-run effects of spending shocks. (iii) The degree of openness, measured here by the share of imports over GDP, does not seem to be an important driving force of spending multipliers. With respect to compositional effects, (iv) a higher share of government investment in total spending has a positive effect on the size of short-run multipliers, whereas (v) a larger wage component of government spending is associated with smaller short-run multipliers.

2.7 Conclusion

This chapter has estimated vector autoregressions with drifting coefficients and stochastic volatility for the euro area, with the aim of investigating changes in the effects of government spending shocks over the period 1980-2008 and, based on second-stage inference, revealing the driving forces of the fiscal transmission mechanism.

Our results indicate that the effectiveness of spending shocks in stimulating economic activity has decreased over time. The estimated short-run multipliers are highest in the late 1980s when they reached values above unity, but they fall afterwards to values closer to 0.5 in the current decade. Longer-term multipliers show a more than two-fold decline since the 1980s. These results suggest that other components of aggregate demand are increasingly being crowded out by spending-based fiscal expansions. In particular, the response of private consumption to government spending shocks has become substantially weaker over time. We also document a weaker response of real wages, whereas the nominal interest rate shows a stronger reaction.

With respect to the driving forces of time variation, our evidence points towards availability of credit as one of the main determinants of the short-run effects of government spending. Furthermore, a lower share of government investment and a larger wage component in total spending seem to have contributed to the observed decline in short-run multipliers. Finally, our results suggest that rising government debt is associated with declining spending multipliers at longer horizons, and thus increasingly
negative longer-term consequences of fiscal expansions.

2.A Details of the Gibbs sampler

This appendix outlines the details of the Gibbs sampling algorithm used for estimation of the TVP-VAR model. The algorithm generates a Markov chain which is a sample from the joint posterior distribution of the VAR parameters (i.e. coefficient states, covariance states, volatility states, and hyperparameters). It combines elements of Benati and Mumtaz (2007), Cogley and Sargent (2005), and Primiceri (2005), with a few additional restrictions on the structure of the hyperparameters.

In the following, \( x^t \) denotes the history of \( x \) up to time \( t \), i.e. \( x^t = [x_1', x_2', x_3', \ldots, x_t']' \), and \( T \) denotes the sample length. Furthermore, re-write the observation equation (2.2) in the main text conveniently as

\[
y_t = X_t' \beta_t + u_t, \tag{2.6}
\]

where \( X_t' = I \otimes [y_{t-1}', y_{t-2}', y_{t-3}', \ldots, y_{t-p}', z_t'] \). The estimation of the model proceeds in the following four steps.

(i) Drawing coefficient states \( \beta^T \). Conditional on \( A^T \) and \( H^T \), one obtains a history \( R^T \). Then, conditional on \( y^T, R^T \), and \( Q \), the observation equation (2) is linear with Gaussian innovations and a known covariance matrix. The posterior density of the coefficients can be factored as\(^{26}\)

\[
f(\beta^T|y^T, R^T, Q) = f(\beta_T|y^T, R^T, Q) \prod_{t=1}^{T-1} f(\beta_t|\beta_{t+1}, y^t, R_t, Q), \tag{2.7}
\]

where

\[
\beta_t|\beta_{t+1}, y^t, R^T, Q \sim N(\beta_{t|t+1}, P_{t|t+1}),
\]

\[
\beta_{t|t+1} = E[\beta_t|\beta_{t+1}, y^t, R^T, Q],
\]

\[
P_{t|t+1} = E[P_t|P_{t+1}, y^t, R^T, Q].
\]

\(^{26}\)Conditioning factors which are redundant in the respective step are omitted.
The conditional means and variances can be computed using the Kalman filter and a backward recursion (see Carter and Kohn, 1994). The Kalman filter delivers

\[ P_{t|t-1} = P_{t-1|t-1} + Q, \quad K_t = P_{t|t-1}X_t'P_{t|t-1}^{-1}X_t + R_t, \]

\[ \beta_{t|t} = \beta_{t-1|t-1} + K_t(y_t - X_t'\beta_{t-1|t-1}), \quad P_{t|t} = P_{t|t-1} - K_tX_t'P_{t|t-1}. \]

The initial values \( \beta_{0|0} \) for this recursion are the OLS point estimates from the initial sample, and the initial value \( P_{0|0} \) is their covariance matrix. The initial \( R_t \) is the OLS covariance matrix of the reduced-form VAR model. The covariance matrix \( Q \) is a scaled version of the variance-covariance matrix of the coefficients. The Kalman filter delivers as its last points \( \beta_{T|T} \) and \( P_{T|T} \). Draws from (2.7) are then obtained by a backward recursion. The first point in the backward recursion is a draw from \( N(\beta_{T|T}, P_{T|T}) \). The remaining draws are from \( N(\beta_{t|t+1}, P_{t|t+1}) \), where the means and variances are derived as follows:

\[ \beta_{t|t+1} = \beta_{t|t} + P_{t|t}P_{t+1|t}^{-1}(\beta_{t+1} - \beta_{t|t}), \quad P_{t|t+1} = P_{t|t} - P_{t|t}P_{t+1|t}^{-1}P_{t|t}. \]

(ii) **Drawing covariance states** \( A^T \). Conditional on \( y^T, \beta^T, \) and \( H^T \), the system of equations (2.6) can be written as follows:

\[ A_t(y_t - X_t'\beta_t) = A_t\hat{y}_t = H_t^{1/2}v_t. \] (2.8)

Moreover, \( A_t \) is lower diagonal (with ones on the main diagonal) such that (2.8) can be re-written as

\[ \hat{y}_t = Z_t\alpha_t + H_t^{1/2}v_t, \] (2.9)

where \( \alpha_t \) is defined as in the main text and \( Z_t \) has the structure

\[
Z_t = \begin{bmatrix}
0 & \cdots & \cdots & 0 \\
-\hat{y}_{1,t} & 0 & \cdots & : \\
0 & (-\hat{y}_{1,t}, -\hat{y}_{2,t}) & \cdots & : \\
: & \cdots & \cdots & 0 \\
0 & \cdots & 0 & (-\hat{y}_{1,t}, \cdots, -\hat{y}_{n-1,t})
\end{bmatrix},
\]
where $n$ denotes the number of variables in the VAR model. The system of equations (2.9) has a Gaussian but non-linear state-space form. However, under the assumption of (block) diagonality of $S$ the problem becomes linear (see Primiceri, 2005). The forward (Kalman filter) and backward recursions of the previous step can then be applied equation by equation. Hence, the procedure allows to recover $\alpha^T$ through

$$
\alpha_{i,t|t+1} = E[\alpha_{i,t} | \alpha_{i,t+1}, y^t, \beta^T, H^T, S_i],
$$

$$
\Lambda_{i,t|t+1} = \text{var}[\alpha_{i,t} | \alpha_{i,t+1}, y^t, \beta^T, H^T, S_i],
$$

where $\alpha_{i,t}$ is the block of $\alpha_t$ corresponding to the $i$-th equation and $S_i$ is the associated $i$-th block of $S$. The initial values for the Kalman filter are obtained from a decomposition of the OLS covariance matrix.

(iii) Drawing volatility states $H^T$. To sample the stochastic volatilities, the univariate algorithm of Jacquier, Polson, and Rossi (1994) is applied to each element of $H_t$. The orthogonalized residuals $v_t = A_t u_t$ are observable conditional on $y^T$, $\beta^T$, and $A^T$. We can use the univariate setting since the stochastic volatilities are assumed to be independent, following Cogley and Sargent (2005). Jacquier, Polson, and Rossi (1994) show that the conditional kernel is

$$
f(h_{i,t}|h_{-i,t-1}, v^T_i, w_i) \propto f(h_{i,t}|h_{i,t-1}, h_{i,t+1}, v^T_i, w_i),
$$

where $w_i$ is the $i$-th diagonal element of $W$ and $h_{-i,t}$ represents the vector of $h$’s at all other dates. Using Bayes’ theorem, the conditional kernel can be expressed as

$$
f(h_{i,t}|h_{i,t-1}, h_{i,t+1}, v^T_i, w_i) \propto f(u_{i,t}|h_{i,t}) f(h_{i,t}|h_{i,t-1}) f(h_{i,t+1}|h_{i,t})
\propto h_{i,t}^{-1.5} \exp\left(-\frac{v^2_{i,t}}{2h_{i,t}}\right) \exp\left(-\frac{(\ln h_{i,t} - \mu_{i,t})^2}{2\sigma_{ic}^2}\right),
$$

where $\mu_{i,t}$ and $\sigma_{ic}^2$ are the conditional mean and the conditional variance of $h_{i,t}$ implied by equation (2.5) in the main text and knowledge of $h_{i,t-1}$ and $h_{i,t+1}$. For a geometric
random walk these parameters are

\[ \mu_{i,t} = 0.5(\log h_{i,t-1} + \log h_{i,t+1}) \quad \text{and} \quad \sigma^2_{ic} = 0.5w_i. \]

In practice \( h_{i,t+1} \) is taken from the previous Gibbs iteration.\(^{27}\) Jacquier, Polson, and Rossi (1994) propose a Metropolis step instead of a Gibbs step, because the normalizing constant is expensive to calculate in (2.10). Hence, one draws from a stand-in density and then uses the conditional likelihood \( f(u_{i,t}|h_{i,t}) \) to calculate the acceptance probability for that draw. Cogley and Sargent (2005) suggest to use the log-normal density implied by equation (2.5) in the main text as the stand-in density:

\[ g(h_{i,t}) \propto h_{i,t}^{-1}\exp \left( -\frac{(\log h_{i,t} - \mu_{i,t})^2}{2\sigma^2_{ic}} \right). \]

The acceptance probability for the \( m \)-th draw is

\[ q_m = \frac{f(v_{i,t}|h_{i,t}^m)g(h_{i,t}^m)}{g(h_{i,t}^m)} \frac{g(h_{i,t}^{m-1})}{f(v_{i,t}|h_{i,t}^{m-1})} g(h_{i,t}^{m-1}) = \frac{(h_{i,t}^m)^{-1/2} \exp \left( -v_{i,t}^2/2h_{i,t}^m \right)}{(h_{i,t}^{m-1})^{-1/2} \exp \left( -v_{i,t}^2/2h_{i,t}^{m-1} \right)}, \]

where \( h_{i,t}^m = h_{i,t}^{m-1} \) if the draw is rejected. This algorithm is applied on a date-by-date basis to each element of \( u_t \). The formulas are slightly different for the first and last element. For the first element we have

\[ \mu_{i1} = \sigma^2_{ic} \left( \frac{\mu_{i0}}{\sigma^2_{h0}} + \frac{\log h_{i,t+1}}{w_i} \right) \quad \text{and} \quad \sigma^2_{ic} = \frac{\sigma^2_{h0}w_i}{\sigma^2_{h0} + w_i}, \]

and the acceptance probability is equal to one since there is no previous draw. For the last element we have

\[ \mu_{iT} = \log h_{i,t-1} \quad \text{and} \quad \sigma^2_{ic} = w_i, \]

where the prior on the distribution of \( \log h_0 \), providing values for the mean \( \mu_{i0} \) and the variance \( \sigma^2_{h0} \), is described in Appendix 2.B.

\(^{27}\)In the first iteration, the squared orthogonalized residuals \( v_{i,t}^2 \) are used to initialize the volatilities, which are calculated by applying the OLS estimates from the initial sample on the actual sample.
(iv) **Drawing hyperparameters.** The hyperparameters of the model are the covariance matrices of the innovations, i.e. $Q$ (coefficient states), $S$ (covariance states), and $W$ (volatility states). Conditional on $y^T$, $\beta^T$, $A^T$, and $H^T$, the state innovations are observable. Since the hyperparameters are assumed to be independent, each covariance matrix can be drawn from its respective distribution. Since we have restricted the hyperparameter matrix $Q$ to be diagonal, its diagonal elements $q_i$ have univariate inverse Gamma distributions with scale parameter $\gamma_{q,0}^{q}$ and degrees of freedom $\delta_{q,0}^{q}$:

$$f(q_i | y^T, \beta^T) = IG\left(\frac{\gamma_{q,1}^{q}}{2}, \frac{\delta_{q,1}^{q}}{2}\right),$$

where $\delta_{q,1}^{q} = \delta_{q,0}^{q} + T$ and $\gamma_{q,1}^{q} = \gamma_{q,0}^{q} + \sum_{t=1}^{T} \epsilon_{t,t}^2$ (see e.g. Kim and Nelson, 1999). Similarly, restricting $S$ to be diagonal, each of its diagonal elements $s_i$ has an inverse Gamma distribution with scale parameter $\gamma_{s,0}^{s}$ and degrees of freedom $\delta_{s,0}^{s}$:

$$f(s_i | y^T, A^T) = IG\left(\frac{\gamma_{s,1}^{s}}{2}, \frac{\delta_{s,1}^{s}}{2}\right),$$

where $\delta_{s,1}^{s} = \delta_{s,0}^{s} + T$ and $\gamma_{s,1}^{s} = \gamma_{s,0}^{s} + \sum_{t=1}^{T} \nu_{t,t}^2$. Finally, the diagonal elements $w_i$ of $W$ have univariate inverse Gamma distributions with scale parameter $\gamma_{w,0}^{w}$ and degrees of freedom $\delta_{w,0}^{w}$:

$$f(w_i | y^T, H^T) = IG\left(\frac{\gamma_{w,1}^{w}}{2}, \frac{\delta_{w,1}^{w}}{2}\right),$$

where $\delta_{w,1}^{w} = \delta_{w,0}^{w} + T$ and $\gamma_{w,1}^{w} = \gamma_{w,0}^{w} + \sum_{t=1}^{T} \omega_{t,t}^2$.

**Summary.** The Gibbs sampling algorithm is summarized as follows:

1. Initialize $R^T$, $Q$, $S$, and $W$.
2. Draw coefficients $\beta^T$ from $f(\beta^T | y^T, R^T, Q)$.
3. Draw covariances $A^T$ from $f(A^T | y^T, H^T, S)$.
4. Draw volatilities $H^T$ from $f(H^T | y^T, \beta^T, A^T, W)$.
5. Draw hyperparameters from $f(q_i | y^T, \beta^T)$, $f(s_i | y^T, A^T)$, and $f(w_i | y^T, H^T)$.
6. Go to step 2.
2.B Calibration of the priors

This appendix discusses the calibration of our priors. We closely follow common choices in the TVP-VAR literature and impose relatively conservative priors, particularly on the hyperparameters (see e.g. Benati and Mumtaz, 2007; Cogley and Sargent, 2002, 2005; Primiceri, 2005).

However, unlike most previous studies those priors are not calibrated based on OLS estimates from an initial training sample which is then discarded. This strategy would force us to sacrifice part of our already relatively short sample. Instead, we calibrate our priors based on OLS estimates from the full sample. This type of strategy is suggested by Canova (2007) and Canova and Ciccarelli (2009) for cases where a training sample is not available. A fixed-coefficient VAR model is thus estimated by OLS (equation by equation) on the full sample from 1980Q1 to 2008Q4.

**VAR coefficients.** Let \( \hat{\beta} \) denote the OLS estimate of the VAR coefficients and \( \hat{\Xi} \) their covariance matrix. We set

\[
\beta_0 \sim N(\hat{\beta}, 4 \times \hat{\Xi}),
\]

where the variance scaling factor increases the uncertainty about the size of the VAR coefficients in the initial sample versus the actual sample.

**Elements of \( H_t \).** Denote the OLS estimate of the VAR covariance matrix as \( \hat{\Sigma} \). We apply a triangular decomposition of this matrix similar to equation (2.4) in the main text, \( \hat{\Sigma} = \hat{\Psi}^{-1}\hat{\Phi}(\hat{\Psi}^{-1})' \), and denote the vector of diagonal elements of \( \hat{\Phi} \) as \( \phi_0 \). Our prior for the diagonal elements of the matrix \( H_t \) is

\[
h_0 \sim N(\phi_0, 10 \times I).
\]

The variance scaling factor 10 is arbitrary but large relative to the mean \( \phi_0 \).
Elements of $A_t$. Denote the vector of non-zero off-diagonal elements of $\hat{\Psi}$ as $\psi_0$, ordered by rows. The prior for the elements of $A_t$ is

$$\alpha_0 \sim N(\psi_0, 10 \times \text{diag}(\psi_0)),$$

where the variance of $\alpha_0$ is scaled up taking into the magnitude of the respective elements of the mean $\psi_0$, as in Benati and Mumtaz (2007).

Hyperparameters. The prior on the diagonal elements of the coefficient state error variance $Q$ is also of the inverse Gamma type:

$$q_i \sim IG\left(\frac{\gamma_{q,i}^q}{2}, \frac{\delta_{q,0}^q}{2}\right),$$

where $\gamma_{q,i}^q = k_Q \times \hat{\xi}_i$, where $\hat{\xi}_i$ denotes the $i$-th diagonal element of the OLS covariance matrix $\hat{\Xi}$ and $k_Q = 10^{-4}$. Hence, our prior attributes only 0.01 percent of the uncertainty surrounding the OLS estimates to time variation, following Cogley and Sargent (2002). The degrees of freedom $\delta_{q,0}^q$ are set to one, which is the minimum for the prior to be proper. We thus put as little weight on the prior as possible. The prior on the diagonal elements of the hyperparameter matrix $S$ for the covariance states is also of the inverse Gamma type:

$$s_i \sim IG\left(\frac{\gamma_{s,i}^s}{2}, \frac{\delta_{s,0}^s}{2}\right),$$

where $\gamma_{s,i}^s = k_S \times \hat{\psi}_i$, where $\hat{\psi}_i$ denotes the $i$-th diagonal element of the OLS covariance matrix $\hat{\Psi}$ and $k_S = 10^{-2}$. Here we follow Primiceri (2005), who makes similar choices for a block diagonal structure of $S$. The degrees of freedom $\delta_{s,0}^s$ are again set to the minimum value of one. The prior on the diagonal elements of the variance $W$ for the volatility states is also of the inverse Gamma type:

$$w_i \sim IG\left(\frac{\gamma_{w,i}^w}{2}, \frac{\delta_{w,0}^w}{2}\right),$$

where $\gamma_{w,i}^w = k_W$. We set $k_W = 10^{-4}$ and $\delta_{w,0}^w = 1$. The parameters of the distribution are the same as in Cogley and Sargent (2005), and Benati and Mumtaz (2007).
2.C Convergence of the Markov chain

This appendix assesses the convergence of the Markov chain produced by the Gibbs sampler. We apply three types of convergence checks to the VAR coefficients, the covariances, and the volatilities. The hyperparameters are omitted in these checks, because they are not the direct objects of interest.

The first convergence check are the diagnostics due to Raftery and Lewis (1992), which are used to assess the total number of iterations required to achieve a certain precision, and the minimum burn-in period and thinning factor. The parameters for the diagnostics are specified as follows: quantile = 0.025; desired accuracy = 0.025; required probability of attaining the required accuracy = 0.95. We generate a Markov chain with 5,000 draws which is then used as the input for the diagnostics as suggested by Raftery and Lewis (1992). Table 2.4 reports the diagnostics. For all three state vectors, the required number of runs is far below the total number of iterations actually applied. The same holds for the number of burn-in replications and the thinning factor. The choices made to generate the Markov chain therefore seem to be validated.

The second convergence check are the inefficiency factors (IFs) for the posterior estimates of the parameters. The IF is the inverse of Geweke’s (1989) relative numerical efficiency measure, i.e. \( IF = 1 + 2 \sum_{k=1}^{\infty} \rho_k \), where \( \rho_k \) is the k-th order autocorrelation of the chain. This diagnostic therefore serves to judge how well the chain mixes. Primiceri (2005) argues that low autocorrelations suggest that the draws are close to independent, which increases the efficiency of the algorithm. We use a 4 percent tapered window for the estimation of the spectral density at frequency zero. Values of the IFs below or around 20 are regarded as satisfactory, according to Primiceri (2005). The left panels of Figure 2.14 report the IFs for the state vectors. The IFs are far below 20 for the coefficients and the covariances, but around 30 to 35 for the volatilities. Compared to the results reported e.g. in Primiceri (2005) and considering the lower number of observations in our sample, however, these results still seem satisfactory.

The final convergence test applied is the convergence diagnostic (CD) due to Geweke (1992). According to Koop (2003), this diagnostic is based on the idea that,

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28See Koop (2003), Chapter 4, for a review of convergence diagnostics.
Table 2.4: Raftery and Lewis (1992) diagnostics\textsuperscript{a,b}

<table>
<thead>
<tr>
<th>Estim. parameters</th>
<th>Thinning factor</th>
<th>Burn-in replic.</th>
<th>Total runs</th>
</tr>
</thead>
<tbody>
<tr>
<td>Coefficients</td>
<td>4068</td>
<td>1</td>
<td>2</td>
</tr>
<tr>
<td>Covariances</td>
<td>452</td>
<td>1</td>
<td>10</td>
</tr>
<tr>
<td>Volatilities</td>
<td>678</td>
<td>1</td>
<td>4</td>
</tr>
</tbody>
</table>

\textsuperscript{a} The parameters for the Raftery and Lewis (1992) diagnostics are as follows: quantile = 0.025, desired accuracy = 0.025, required probability of attaining the required accuracy = 0.95.

\textsuperscript{b} The results are based on 5,000 iterations of the Gibbs sampler with zero burn-in replications and thinning factor equal to one.

if a sufficiently large number of draws has been taken, the posterior estimates based on the first half of draws should be essentially the same as the estimates based on the second half of draws. If they are very different, either too few draws have been taken and estimates are inaccurate or the effects of the initial values of the chain have not worn off. We therefore divide the 1,000 draws from the posterior distribution into a first set of $N_1 = 100$ draws, a middle set of 500 draws, and a last set of $N_2 = 400$ draws, as suggested by Koop (2003). The middle set of draws is dropped to make it likely that the first set and the last set are independent of each other, which is assessed by the diagnostic. The convergence diagnostic is given by

$$CD = \frac{\hat{\theta}_1 - \hat{\theta}_2}{\hat{\sigma}_1/\sqrt{N_1} + \hat{\sigma}_2/\sqrt{N_2}} \rightarrow N(0,1),$$

by a central limit theorem, where $\hat{\theta}_i$ and $\hat{\sigma}_i/\sqrt{N_i}$ denote the posterior means of the parameters and their numerical standard errors based on the $i$-th set of draws, for $i = 1, 2$ (see Koop, 2003). We plot the $p$-values for the null hypothesis that the two sets of draws are the same in the right panels of Figure 2.14. The $p$-values are mostly larger than conventional significance levels for the VAR coefficients and the covariances, indicating that a sufficiently large number of draws has been taken for these parameters. The null hypothesis is often rejected for the volatilities, but this outcome did not change when a larger number of draws was taken.

To summarize, the coefficients and covariances have in general better convergence properties than the volatilities. Since the focus of our analysis is on impulse responses
which are mainly determined by the contemporaneous relations among variables and the VAR coefficients rather than the size of stochastic shocks, we conclude that the convergence properties of the Markov chain are overall satisfactory.

2.D Detailed data description

This appendix provides details on the data definitions used in the main text. Throughout, AWM refers to the Area-Wide Model database (see Fagan et al., 2005), BIS to the Bank of International Settlements macro-economic series, CE to the Consensus Eco-
Chapter 2

nomic survey data, EC to the European Commission forecasts and PPP to the data set provided by Paredes et al. (2009), to which we refer for details on the construction of the fiscal variables. All quarterly series are provided in seasonally adjusted terms from the original sources, except for the HICP of which we take annual differences.

- **Government spending**: Sum of nominal general government final consumption expenditure (variable GCN in PPP) and nominal general government investment (variable GIN in PPP), euro area aggregates, scaled by GDP deflator plus labor force and transformed into natural logarithms.

- **GDP**: Aggregate euro area real gross domestic product, variable YER in the AWM database, where it is calculated as a weighted average of national variables. The original source of GDP and its components in AWM is Eurostat; the variables are then re-scaled to the ECU-euro corrected level of 1995 and backdated with rates of growth of the original AWM series.29

- **Private consumption**: Aggregate euro area private consumption, constructed by multiplying real private consumption (variable PCR in AWM) with the private consumption deflator (variable PCD in AWM), divided by GDP deflator plus labor force and transformed into natural logarithms.

- **Interest rate**: Weighted euro area short-term nominal interest rate, variable STN in AWM, where it is calculated as a weighted average of national variables taken from the ECB Monthly Bulletin and backdated with the corresponding series contained in the original database (source: Bank of International Settlements and European Commission’s AMECO database).

- **Private investment**: Aggregate euro area total economy gross investment minus general government investment (nominals), scaled by GDP deflator plus labor force and transformed into natural logarithms. Total economy investment corresponds to the variable ITR in AWM, government investment is the variable GIN in PPP.

29The weights used in AWM are based on constant GDP at market prices for the euro area for 1995.
• **Wage rate:** Nominal hourly wage per head (variable WRN in AWM) divided by GDP deflator. The nominal wage in AWM is calculated as a weighted average of national variables.

• **Net taxes:** Non-interest nominal general government revenue (variable TOR in PPP) minus transfers, which include all expenditure items except government consumption, government investment, and interest payments (variable INP in PPP), scaled by GDP deflator plus labor force and transformed into natural logarithms. The general government primary balance is thus the difference between net taxes and government spending.

• **Inflation rate:** Annual rate of change of the Harmonized Index of Consumer Prices, i.e. variable HICP in AWM, where it is calculated as a weighted average of national variables using 1995 HICP weights.

• **GDP deflator:** Index with base year 1995, variable YED in AWM. Deflators in AWM are taken directly from the corresponding ECB Monthly Bulletin series, which are compiled by ECB staff as a weighted average of the national deflators using purchasing power parity adjusted weights.

• **Labor force:** Total euro area labor force, persons, variable LFN in AWM.

• **Debt-to-GDP ratio:** Ratio of the outstanding (end-of-period) aggregate euro area stock of nominal public debt over nominal annual euro area GDP, variable GDN.YEN in AWM.

• **Imports-to-GDP ratio:** Ratio of nominal quarterly aggregate euro area imports (variable MTR times MTD in AWM) over nominal quarterly euro area GDP.

• **Credit to households over GDP:** Outstanding total euro area (end-of-period) stock of bank loans to households, variable BIS.M.Q.COVA.XM.03 in BIS, divided by the sum of nominal euro area GDP of the last four consecutive quarters.

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30 The labor force is used as a proxy for total population, since quarterly data on total population is not available from AWM for the entire sample period.
• **Government budget deficit-to-GDP ratio, forecast (EC):** (Minus) general government balance as percentage of GDP, one-year ahead forecasts by EC published in November of the previous year. Forecasts for the euro area are available from 1999 onwards; for previous years up to 1982, aggregate forecasts are constructed from forecasts for Belgium, France, Germany, Greece, Ireland, Italy, Luxembourg, and the Netherlands by aggregating the individual country series using as weights constant GDP at market prices for 1995.

• **GDP growth, forecast (EC):** Annual real GDP growth rate, one-year ahead forecasts by EC published in November of the previous year. Forecasts for the euro area are only available from 1999 onwards; for previous years up to 1982, aggregate forecasts are constructed as described above.

• **Government budget deficit growth, forecast (CE):** Consensus mean forecast of (minus) the general government budget balance, converted into growth rates, minus the Consensus mean forecast of consumer price inflation. Both forecasts are computed as the average of one-year ahead forecasts made in each month of the previous year. Forecasts for the euro area are available from 2003 onwards; for previous years up to 1994, aggregate forecasts are constructed from forecasts for France, Germany, and Italy by aggregating the individual country series using as weights constant GDP at market prices for 1995.

• **GDP growth, forecast (CE):** Consensus mean forecast of the annual real GDP growth rate, computed as the average of one-year ahead forecasts made in each month of the previous year. Forecasts for the euro area are available from 2003 onwards; for previous years up to 1992, aggregate forecasts are constructed as described above.

• **Consumption growth, forecast (CE):** Consensus mean forecast of the annual real private consumption growth rate, computed as described above.

• **Interest rate, forecast (CE):** Consensus mean forecast of the short-term (3-month) nominal interest rate, computed as described above.
Chapter 3

Expectations-Based Identification of Government Spending Shocks under Policy Foresight

Abstract

This chapter is concerned with the econometric problems of structural vector autoregressive analysis of the effects of government spending that are created by the presence of news or foresight about future fiscal policy. Using a combination of theory and stochastic simulations, the chapter investigates whether incorporating data that captures expectations on spending in VAR models is useful to address (i) the non-fundamentalness problem due to policy foresight and (ii) the problem of identifying structural spending shocks in the presence of policy foresight. In particular, the chapter demonstrates under which conditions and how expectations-based identifying restrictions can be applied to distinguish between anticipated and unanticipated shocks. In an application to U.S. data, the expectations-based approach that is proposed based on this analysis indicates a weaker impact of government spending on output, consumption, and investment than the most commonly applied fiscal VAR approach, which does not take into account the possibility of policy foresight.

3.1 Introduction

Discretionary changes in government spending can sometimes be anticipated in advance of their implementation, for example when spending programs are pre-announced in political speeches and news statements or when armed conflicts create expectations
of military spending hikes. Such phenomena of policy foresight pose the following two challenges to attempts of quantifying the macroeconomic effects of government spending by empirical structural VAR (SVAR) methods.

First, if economic agents have access to information on fundamentals that affect future government spending but that are not yet incorporated in the actual data on government spending, then economic equilibria can have non-fundamental moving average representations. This means that the fundamental structural spending shocks cannot be recovered from present and past data on government spending by SVAR methods. As a consequence, ignoring the influence of policy foresight is likely to lead to biased estimates, as shown by Leeper, Walker, and Yang (2011), Ramey (2011b), and Yang (2005). Second, even if these non-fundamentalness issues could be circumvented, the structural shocks still need to be recovered from reduced-form VAR estimates by an appropriate identification strategy. This strategy not only needs to distinguish spending shocks from other economic shocks, but it also needs to distinguish between anticipated and unanticipated spending shocks. Good identifying restrictions are therefore especially hard to come by under policy foresight.

This chapter investigates whether incorporating expectations survey data in a vector autoregression can be useful to address these econometric difficulties. A number of recent studies indeed suggest to exploit information in forward-looking variables in fiscal VARs (e.g. Auerbach and Gorodnichenko, 2010; Ramey, 2011b; Sims, 2009) and some of those studies support this idea by simulation results. Laubach (2008) has also proposed to use direct measures of expectations on fiscal variables to address the econometric challenges posed by anticipation effects. To obtain a deeper understanding of such proposals and to make a precise link to the work of Leeper, Walker, and Yang (2011) and related work by Mertens and Ravn (2010), this chapter provides analytical results based on a standard growth model and tightly connected simulation evidence. In particular, the chapter studies the properties of identifying restrictions on government spending shocks that are based on model-consistent expectations. The candidate identification strategy is as follows: (i) an unanticipated or surprise spending increase is identified as an increase in the expectational error between today’s spending and expectations thereof formed yesterday, while (ii) an anticipated spending increase
is identified by an increase in expected spending tomorrow that is orthogonal to this expectational error.

The chapter first derives conditions under which incorporating expectations in a VAR model works to solve the non-fundamentalness problem due to policy foresight. For standard information flows, the requirement that is found to be sufficient is that expectations on future spending up to the anticipation horizon of economic agents should be included as an endogenous variable in the regression. Then, regarding the identification problem, the chapter shows that the success of the above identification approach depends on the type of endogenous reactions of government expenditures to the state of the economy, if one allows for the possibility that spending is not entirely exogenously determined. Surprise spending shocks are robustly identified if government spending reacts with some lag to other economic shocks, which is also assumed under the short-run restrictions of the standard recursive fiscal SVAR approach (see Blanchard and Perotti, 2002; Fatás and Mihov, 2001). However, the expectations-based approach may fail to correctly recover the effects of anticipated spending shocks even if government spending reacts to other shocks with a lag. The effects of anticipated shocks can only be correctly identified if all other relevant exogenous shocks are known, observed, and conditioned upon; that is, under fairly restrictive requirements. The expectations-based approach is thus found to be useful for the identification of surprise spending shocks when it is combined with standard short-run exogeneity restrictions.

Given these findings, the chapter focuses on the effects of surprise spending shocks in an application of the approach to the U.S., where data on expected federal government spending from the Survey of Professional Forecasters is used to measure expectations, presuming that this data does capture the expectations of economic agents (or market participants). This exercise reveals important differences in the effects of government spending according to the expectations-based approach in comparison to the standard fiscal SVAR approach. According to the expectations-based approach, an unexpected spending increase has positive short-run effects on output but negative effects at longer horizons, going along with pronounced declines in private consumption and investment. The standard SVAR approach, on the other hand, predicts an increase in consumption and investment in the short run and larger multipliers on GDP at longer horizons.
Following Ramey (2011b), the chapter also shows that the standard SVAR shocks are Granger-caused by the survey expectations (i.e. they are predictable) whereas the surprise shocks identified on the basis of expectational errors are indeed truly unpredictable. Taking into account policy foresight thus has important implications for the empirical effects of spending shocks.

The fact that anticipated shocks are not identified is of course a strong limitation if those shocks are the only relevant shocks to government spending or if they fall on different sub-items of total spending than unanticipated shocks. Surprise spending shocks are indeed more likely to fall on consumption expenditures under the standard definition of government spending, i.e. government consumption plus investment expenditures (see e.g. Blanchard and Perotti, 2002), because implementation lags are usually argued to be more relevant for government investment (see e.g. Leeper, Walker, and Yang, 2010). The empirical results in this chapter should be interpreted with these possible limitations in mind, despite the reassuring fact that some of the results reported below suggest that surprise spending shocks have been important driving forces of U.S. government spending in the past.1

Several other recent studies have addressed the econometric issues created by foresight on government spending. Ramey (2011b) applies a narrative approach which exploits military spending episodes, newspaper sources, and forecast errors based on survey data. However, although she also uses expectational errors to identify surprise spending shocks, Ramey does not add expectations on future spending in the VAR, which does not lead to a fundamental moving average representation. Fisher and Peters (2010) identify spending shocks by innovations to excess stock returns of military contractors, an approach which is applicable to defense-related expenditures only. Furthermore, Mertens and Ravn (2010) propose an SVAR estimator for permanent spending shocks based on Blaschke matrices, following Lippi and Reichlin (1994), which can be applied when the identifying assumptions pin down the Blaschke factor. Kriwoluzky (2010) directly estimates a vector moving average model, with model-based identifying restrictions on anticipated spending shocks. Forni and Gambetti (2010) es-

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1See, in particular, the variation in the expectational errors in Figures 3.16, the tightness of the uncertainty bands in Figure 3.17, and the size of the identified shocks shown in Figure 3.20.
timate a large factor model which is arguably not affected by the non-fundamentalness problem (see Forni, Giannone, Lippi, and Reichlin, 2009), identifying a spending shock by various sign restrictions. However, the latter three approaches all rely on the correct specification of the identifying theoretical model. An expectations-based approach has the advantage of applying seemingly less restrictive identifying assumptions.

The remainder of the chapter is structured as follows. The next section explores the econometric problems created by foresight in a standard growth model and studies the usefulness of an expectations-based approach in addressing those problems. Section 3.3 provides connected simulation evidence. Section 3.4 analyzes the robustness of the approach to, not exclusively, alternative assumptions on the structure of the spending process, and proposes possible adjustments based on this analysis. Section 3.5 discusses the results of the empirical application. Section 3.6 concludes.

3.2 Policy foresight: problems and solutions

This section explores the problems induced by foresight on government spending in a simple analytical example, following Leeper, Walker, and Yang (2011) and Mertens and Ravn (2010). Leeper, Walker, and Yang (2011) analyze the econometric implications of foresight on future tax rates. Mertens and Ravn (2010) focus on government spending, in order to derive an SVAR estimator which is applicable in the face of permanent spending shocks. This section discusses potential solutions when economic agents’ expectations can be observed. Throughout, the data is assumed to be generated by a version of the simple neoclassical growth model due to Hansen (1985). The main results would also hold in a larger DSGE model. The obvious advantage of using a simpler model is that analytical results can be derived more easily.²

3.2.1 Model description

The model economy is inhabited by a continuum of identical, infinitely lived households, whose instantaneous utility depends on consumption \( c_t \) and hours worked \( n_t \). The

²This section focuses on a basic information structure with two periods anticipation and a single news shock, the latter following Mertens and Ravn (2010) and Ramey (2011b). The spending process could also be extended to longer anticipation horizons without affecting the main results.
households provide labor services and physical capital $k_t$ to firms and they pay lump-sum taxes $\tau_t$ to the government. Time is indexed by $t = 0, 1, 2, \ldots, \infty$. All variables are denoted in real terms. The objective of a representative household in period $t$ is to maximize expected discounted utility

$$ E_t \sum_{s=0}^{\infty} \beta^s (\log c_{t+s} - A n_{t+s}), \quad \beta \in (0, 1), \quad A > 0, $$

where $E_t$ is the mathematical expectations operator conditional on the information available at time $t$. The household’s optimization problem is subject to the period-by-period budget constraint

$$ c_t + i_t + \tau_t = w_t n_t + r_t k_{t-1}, $$

where $w_t$ denotes the hourly wage and $i_t$ denotes investment in physical capital at the rental rate $r_t$. Physical capital accumulates according to the law of motion

$$ k_t = (1 - \delta) k_{t-1} + i_t, \quad \delta \in [0, 1]. \quad (3.1) $$

There is also a continuum of identical, perfectly competitive firms that produce the final consumption good $y_t$. The technology of a representative firm is based on a Cobb-Douglas production function in capital and labor:

$$ y_t = a_t k_{t-1}^\alpha n_t^{1-\alpha}, \quad \alpha \in (0, 1), \quad (3.2) $$

where $a_t$ is total factor productivity (TFP) which is assumed to follow the law of motion

$$ \log a_t = \rho_a \log a_{t-1} + \varepsilon_{a,t}, \quad \rho_a \in [0, 1), \quad (3.3) $$

with $\varepsilon_{a,t} \sim N(0, \sigma_a^2)$. Profit maximization yields the factor prices $w_t = (1 - \alpha) y_t / n_t$ and $r_t = \alpha y_t / k_{t-1}$.

Government spending (i.e. purchases of the final good) $g_t$ is assumed to be financed exclusively by lump-sum taxes, $g_t = \tau_t$, and it is modelled as an exogenous stochastic process:

$$ \log(g_t / \bar{g}) = \rho_g \log(g_{t-1} / \bar{g}) + \varepsilon_{g,t}^u + \varepsilon_{g,t-2}^u, \quad \rho_g \in [0, 1), \quad (3.4) $$
where \( \bar{g} = g \) is the non-stochastic steady state level of government spending, which is taken as given. This process allows for a surprise (unanticipated) shock to government spending \( \varepsilon_{g,t}^u \sim N(0, \sigma_{g,u}^2) \) and a news (anticipated) shock \( \varepsilon_{g,t}^a \sim N(0, \sigma_{g,a}^2) \). When a news shock occurs, the associated change in spending is known by economic agents two periods in advance of its implementation in terms of actual spending.

Combining the household’s budget constraint with the government’s budget constraint and the firm’s first-order conditions, the feasibility constraint reads

\[
c_t + i_t + g_t = y_t. \tag{3.5}
\]

Substituting out the factor prices in the household’s first-order conditions yields the labor/leisure trade-off and the consumption Euler equation:

\[
A c_t = (1 - \alpha) y_t / n_t, \tag{3.6}
\]

\[
c_t^{-1} = \beta E_t R_{t+1} / c_{t+1}, \tag{3.7}
\]

where \( R_t \) denotes the real return on capital, that is

\[
R_t = 1 - \delta + \alpha y_t / k_{t-1}. \tag{3.8}
\]

The rational expectations equilibrium of this model is then the set of sequences \( \{c_t, n_t, i_t, k_t, R_t, y_t, a_t, g_t\}_{t=0}^{\infty} \) satisfying (3.1) to (3.8) and the transversality condition for capital, for given initial values \( k_{-1}, a_{-1}, \) and \( g_{-1}, \) and given sequences of shocks \( \{\varepsilon_{a,t}, \varepsilon_{g,t}^u, \varepsilon_{g,t}^a\}_{t=0}^{\infty}. \)

To obtain an analytical solution to the model, the equilibrium system is log-linearized around the non-stochastic steady state and the log-linearized system is solved by the method of undetermined coefficients (see Uhlig, 1999), similarly as in Mertens and Ravn (2010). In particular, the log-linearized system can be reduced to the following two-dimensional, first-order stochastic difference equation in consumption and capital:

\[
0 = \hat{c}_t - \phi_1 E_t \hat{c}_{t+1} + \phi_2 E_t \hat{a}_{t+1},
\]

\[
0 = \phi_3 \hat{c}_t + \phi_4 \hat{k}_t - \phi_5 \hat{a}_t - \phi_6 \hat{k}_{t-1} + \phi_7 \hat{g}_t,
\]
Table 3.1: Benchmark calibration of the model

<table>
<thead>
<tr>
<th>Parameter</th>
<th>Value</th>
<th>Description</th>
<th>Source/moment*</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\beta$</td>
<td>0.990</td>
<td>Subjective discount factor</td>
<td>KP (1982), Hansen (1985)</td>
</tr>
<tr>
<td>$\alpha$</td>
<td>0.360</td>
<td>Capital share in production</td>
<td>KP (1982), Hansen (1985)</td>
</tr>
<tr>
<td>$\delta$</td>
<td>0.025</td>
<td>Quarterly depreciation rate</td>
<td>KP (1982), Hansen (1985)</td>
</tr>
<tr>
<td>$A$</td>
<td>2.500</td>
<td>Disutility of labor supply</td>
<td>Time for market activities</td>
</tr>
<tr>
<td>$a$</td>
<td>1.000</td>
<td>Steady state TFP</td>
<td>Normalization</td>
</tr>
<tr>
<td>$g/y$</td>
<td>0.080</td>
<td>Government spending over GDP</td>
<td>Federal spending ratio</td>
</tr>
<tr>
<td>$g$</td>
<td>0.100</td>
<td>Government spending</td>
<td>Federal spending ratio</td>
</tr>
<tr>
<td>$\rho_a$</td>
<td>0.950</td>
<td>AR(1) parameter TFP</td>
<td>Hansen (1985)</td>
</tr>
<tr>
<td>$\rho_g$</td>
<td>0.850</td>
<td>AR(1) parameter gov. spending</td>
<td>Gali et al. (2007)</td>
</tr>
<tr>
<td>$\sigma_a$</td>
<td>0.710</td>
<td>Std. dev. TFP shocks (%)</td>
<td>Hansen (1985)</td>
</tr>
<tr>
<td>$\sigma_{g,u}$</td>
<td>1.000</td>
<td>Std. dev. anticip. spend. shocks (%)</td>
<td>Benchmark</td>
</tr>
<tr>
<td>$\sigma_{g,a}$</td>
<td>1.000</td>
<td>Std. dev. unanticip. spend. shocks (%)</td>
<td>Benchmark</td>
</tr>
</tbody>
</table>


given the exogenous processes $\hat{g}_t = \rho_g \hat{g}_{t-1} + \varepsilon^u_{g,t} + \varepsilon^a_{g,t-2}$ and $\hat{a}_t = \rho_a \hat{a}_{t-1} + \varepsilon_{a,t}$, where $\hat{x}_t = \log(x_t/x)$ denotes the log deviation of variable $x_t$ from its steady state value $x$.

The parameters $\phi_i$, $i = 1, 2, 3, \ldots, 7$ are functions of the model parameters and steady state values. The recursive laws of motion describing the solution are as follows:

$$
\hat{k}_t = \eta_{kk} \hat{k}_{t-1} + \eta_{ka} \hat{a}_t + \eta_{kg} \hat{g}_t + \eta_{ke} \varepsilon^a_{g,t-1} + \eta_{ke,1} \varepsilon^u_{g,t}, \quad (3.9)
$$
$$
\hat{c}_t = \eta_{ck} \hat{k}_{t-1} + \eta_{ca} \hat{a}_t + \eta_{cg} \hat{g}_t + \eta_{ce} \varepsilon^a_{g,t-1} + \eta_{ce,1} \varepsilon^u_{g,t}, \quad (3.10)
$$

where the coefficients $\eta$ are non-linear functions of the model parameters and the parameters $\phi$. A detailed derivation is provided in Appendix 3.A. Notice that, according to (3.9) and (3.10), the solution is characterized by the fact that the information set of economic agents in period $t$ includes the shocks $\varepsilon^a_{g,t-1}$ and $\varepsilon^a_{g,t}$ as state variables.3

The model is calibrated in line with the real business cycle literature (see Hansen, 1985; Kydland and Prescott, 1982) and to match selected moments in U.S. quarterly

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3The full information set includes $\{\varepsilon_{a,j}, \varepsilon^u_{g,j}, \varepsilon^a_{g,j}\}_{j=0}^t$. The shocks $\varepsilon_{a,j}$ are incorporated in $\hat{a}_t$ and (up to time $t-1$) in $\hat{k}_{t-1}$. The shocks $\{\varepsilon^u_{g,j}\}_{j=0}^t$ and $\{\varepsilon^a_{g,j}\}_{j=0}^{t-2}$ are also incorporated in $\hat{g}_t$ and $\hat{k}_{t-1}$ whereas $\varepsilon^a_{g,t-1}$ and $\varepsilon^a_{g,t}$ are announced but not yet implemented innovations to spending. Therefore, the latter appear as additional state variables in the recursive laws of motion.
Figure 3.1: Model impulse responses I – both types of spending shocks

Notes. Benchmark calibration; both panels show responses to one percent increases in government spending relative to its steady state value; left panel: surprise spending increase in quarter 0; right panel: news in quarter 0 that spending will increase in quarter 2; responses are measured as relative percentage deviations from steady state.

data over the period 1981Q4 to 2010Q1. The benchmark calibration is provided in Table 3.1. The subjective discount factor $\beta$ is set to 0.99, which implies a steady state annual real interest rate of approximately four percent.\(^4\) The capital share in production $\alpha$ is set to 0.36 and the quarterly depreciation rate $\delta$ is set to 0.025. The parameter $A$ is set to 2.5, which implies that steady state hours worked is close to $1/3$. The steady state ratio of government spending over GDP, $g/y$, is set to its empirical counterpart (the average federal government spending share) of eight percent. Finally, the standard deviation of TFP shocks is set to 0.71 percent and the AR(1) parameter of TFP is set to 0.95, following Hansen (1985). The standard deviations of the two spending innovations are set to one percent and the AR(1) parameter of government spending is set to 0.85 (see e.g. Galí et al., 2007).

Figure 3.1 shows impulse responses to one percent spending shocks of both types. As the model satisfies Ricardian equivalence, a surprise spending increase in quarter 0 (left panel) that is financed by lump-sum taxes has a negative wealth effect on the household’s lifetime income. Consumption declines and, since leisure is a normal good,

\(^4\)The average annual 3-month U.S. treasury bill secondary market rate was approximately equal to five percent over the period 1981Q4 to 2010Q1.
hours worked increase. Although the return on investment increases, the negative investment response is dictated by the feasibility constraint under the chosen calibration. On the other hand, if there is news in quarter 0 that spending will increase in quarter 2 (right panel) the investment response is positive during two quarters and then turns negative. There is an immediate negative wealth effect due to higher future taxes, so consumption declines immediately and hours worked and output increase immediately. The feasibility constraint allows investment to increase during the anticipation period since there is no government absorption of goods and services yet in that period.

### 3.2.2 The non-fundamentalness problem

To characterize the non-fundamentalness problem induced by the news shock $\varepsilon_{g,t}^a$, notice that the coefficient $\eta_{kk}$ is the stable root of the characteristic equation

$$0 = \phi_1\phi_4\eta_{kk}^2 - (\phi_1\phi_6 + \phi_4)\eta_{kk} + \phi_6.$$  

In a unique saddle path solution, this equation has two real roots $\eta_{kk}^+$ and $\eta_{kk}^-$,

$$\eta_{kk}^\pm = (\phi_1^{-1} + \phi_6\phi_4^{-1})/2 \pm \sqrt{(\phi_1^{-1} + \phi_6\phi_4^{-1})^2/4 - \phi_6(\phi_1\phi_4)^{-1}},$$  

(3.11)

only one of which satisfies $|\eta_{kk}| < 1$. Furthermore, it is straightforward to show that the coefficients $\eta_{xx,1}$ and $\eta_{xx,2}$ ($x = k, c$), are related with each other as follows:

$$\eta_{xx,1} = \theta\eta_{xx,2}, \quad \theta = (\phi_1^{-1} + \phi_6\phi_4^{-1} - \eta_{kk})^{-1}.$$  

Inserting the expression for $\theta$ in (3.11), it follows that $|\theta| < 1$. This result implies that, when forming their decisions, economic agents discount more recent news on government spending $\varepsilon_{g,t}^a$ relative to more distant news $\varepsilon_{g,t-1}^a$ at a constant *anticipation rate* given by $\theta$. The reason is that recent news affects spending later than distant news (see Leeper, Walker, and Yang, 2011; Mertens and Ravn, 2010). As noted by

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5To see this, suppose that $|\eta_{kk}^+| = \frac{1}{2}(\phi_1^{-1} + \phi_6\phi_4^{-1}) + \frac{1}{4}(\phi_1^{-1} + \phi_6\phi_4^{-1})^2 - \phi_6(\phi_1\phi_4)^{-1}|^{1/2} < 1$ such that $|\eta_{kk}^-| = \frac{1}{2}(\phi_1^{-1} + \phi_6\phi_4^{-1}) - \frac{1}{4}(\phi_1^{-1} + \phi_6\phi_4^{-1})^2 - \phi_6(\phi_1\phi_4)^{-1}|^{1/2} > 1$. Then $\eta_{kk} = \eta_{kk}^+$ and, by direct calculation, $\theta = (\eta_{kk}^-)^{-1}$, which implies that $|\theta| < 1$. Conversely, if $|\eta_{kk}^+| > 1$ and $|\eta_{kk}^-| < 1$ then $\eta_{kk} = \eta_{kk}^-$ and $\theta = (\eta_{kk}^+)^{-1}$, which again implies that $|\theta| < 1$. 

---
Figure 3.2: Model impulse responses II – unanticipated spending shock

Notes. Changing anticipation rate $\theta$; from thin to thick lines: $\theta$ is changed from 0.58 to 0.93 by changing the discount factor $\beta$ from 0.8 to 0.99; thickest line: benchmark calibration; responses are measured as relative percentage deviations from steady state.

Mertens and Ravn (2010), the result of constant discounting generalizes to other settings (e.g. longer anticipation horizons, more control variables). Mertens and Ravn (2010) also show that the anticipation rate $\theta$ is, inter alia, monotonically increasing in the subjective discount factor $\beta$. This fact is exploited below.

Thus, Figures 3.2 and 3.3 show the impulse responses to the two types of spending shocks when $\beta$ changes from 0.8 to 0.99 from thin to thick lines, implying values for $\theta$ from 0.58 to 0.93. The initial responses of consumption and hours to unanticipated shocks are uniformly stronger for lower discount factors (see Figure 3.2). The reason
is that future utility is then discounted at a higher rate, so households have a lower preference for consumption smoothing. For the same reason, the anticipation rate falls when $\beta$ decreases such that more recent news receives a heavier discount than more distant news. An anticipated increase in government spending thus affects several variables more strongly prior to its implementation in terms of actual spending for lower $\beta$’s and $\theta$’s (see Figure 3.3). For those variables, the differences between the impulse responses to both types of shocks after spending has increased become larger for lower anticipation rates. Compare, for example, the investment responses which are uniformly negative from quarter 2 onwards after the anticipated spending increase but positive for some parameter values after the unanticipated spending increase. The

Notes. News in quarter 0 that spending will increase in quarter 2; see Figure 3.2.
implications of these outcomes are discussed in turn.

The phenomenon of constant discounting is indeed the root of the non-fundamentalness problem. To see this, following Leeper, Walker, and Yang (2011), suppose that an econometrician who is not aware of policy foresight estimates a VAR model in \( \{ \hat{g}_{t-j}, \hat{a}_{t-j}, \hat{k}_{t-j} \}_{j=0}^{\infty} \). According to the equilibrium representation implied by the underlying theoretical model, the econometrician’s observables can be shown to follow the multivariate moving average process

\[
\begin{bmatrix}
\hat{g}_t \\
\hat{a}_t \\
\hat{k}_t 
\end{bmatrix} = \begin{bmatrix}
\frac{1}{1-\rho_z L} & \frac{L^2}{1-\rho_z L} & \frac{0}{1-\rho_z L} \\
0 & \frac{\eta_{kg}L^2 + \eta_{kgz}(1-\rho_z L)(\theta + L)}{(1-\eta_{kg}L)(1-\rho_z L)} & \frac{\frac{\eta_{kg}}{1-\eta_{kg}L}}{(1-\eta_{kg}L)(1-\rho_z L)} \\
\frac{\eta_{ka}}{(1-\eta_{ka}L)(1-\rho_a L)} & \frac{\eta_{ka}L^2 + \eta_{ka}z(1-\rho_a L)(\theta + L)}{(1-\eta_{ka}L)(1-\rho_a L)} & \frac{1}{(1-\eta_{ka}L)(1-\rho_a L)}
\end{bmatrix} \begin{bmatrix}
\varepsilon_{u, g,t} \\
\varepsilon_{u, a,t} \\
\varepsilon_{a,t}
\end{bmatrix},
\]

or

\[ y_t = P(L)\epsilon_t, \quad (3.12) \]

where \( L \) denotes the lag operator, i.e. \( L^s x_t = x_{t-s} \) for \( s \geq 0 \). If the process (3.12) is invertible in non-negative powers of \( L \), then the econometrician can recover the structural shocks as a linear combination of present and past observables, i.e. \( \epsilon_t = P^{-1}(L)y_t \). A necessary and sufficient condition for \( \epsilon_t \) to be fundamental for \( y_t \) is that the zeroes of the determinant of \( P(z) \) do not lie inside the unit circle (see Hansen and Sargent, 1991). The determinant of \( P(z) \) is given by

\[
\det P(z) = \frac{-\eta_{kz, z}(\theta + z)}{(1-\eta_{kz}z)(1-\rho_z z)(1-\rho_a z)},
\]

which has a root inside the unit circle at \( z = -\theta \). The structural shocks \( \{ \varepsilon_{g,j}^u, \varepsilon_{g,j}^a, \varepsilon_{a,j} \}_{j=0}^{\infty} \) thus cannot be recovered from the information set \( \{ \hat{g}_{t-j}, \hat{a}_{t-j}, \hat{k}_{t-j} \}_{j=0}^{\infty} \) since (3.12) is not a fundamental moving average (Wold) representation of the equilibrium time series process. Hence, if the econometrician only observes present and past spending along with present and past TFP and capital, his or her information set is misaligned with the information set of economic agents, who have knowledge of news shocks to future spending over their anticipation horizon already before those shocks have an impact on actual spending.
3.2.3 An expectations-based solution

A natural way to align the two information sets is to incorporate the agents’ expectations on future spending in the econometrician’s information set. Thus, suppose that instead of present and past TFP, the econometrician observes present and past expectations on spending two periods ahead conditional on time \( t \) information. The econometrician thus estimates a VAR in \( \{ \hat{g}_{t-j}, E_{t-j} \hat{g}_{t+2-j}, \hat{k}_{t-j} \}_{j=0}^{\infty} \), where

\[
\begin{bmatrix}
\hat{g}_t \\
E_t \hat{g}_{t+2} \\
\hat{k}_t
\end{bmatrix} = \begin{bmatrix}
\frac{1}{1-\rho_g L} & \frac{L^2}{1-\rho_g L} & 0 \\
\frac{\rho_g^2}{1-\rho_g L} & \frac{1}{1-\rho_g L} & 0 \\
\frac{\eta_k}{(1-\eta_k L)(1-\rho_g L)} & \frac{\eta_k L^2 + \eta_k L (1-\rho_g L)(\theta + L)}{(1-\eta_k L)(1-\rho_g L)} & \frac{\eta_k}{(1-\eta_k L)(1-\rho_a L)}
\end{bmatrix}
\begin{bmatrix}
\varepsilon_{u,t} \\
\varepsilon_{g,t} \\
\varepsilon_{a,t}
\end{bmatrix},
\]

or

\[
y_t^* = P^*(L)\epsilon_t. \tag{3.13}
\]

The determinant of \( P^*(z) \) is given by

\[
\det P^*(z) = \frac{\eta_k a (1 + \rho_g z)}{(1 - \rho_g z)(1 - \eta_k z)(1 - \rho_a z)},
\]

which has one root outside the unit circle at \( z = -\rho_g^{-1} \) and three poles at \( z = \rho_g^{-1}, \eta_k^{-1}, \) and \( z = \rho_a^{-1} \). Hence, (3.13) is indeed an invertible moving average process such that the structural shocks in \( \epsilon_t \) can in principle be recovered from the information set \( \{ \hat{g}_{t-j}, E_{t-j} \hat{g}_{t+2-j}, \hat{k}_{t-j} \}_{j=0}^{\infty} \) by a linear combination of present and past observables. Through the inclusion of economic agents’ expectations on government spending in the econometrician’s information set, (3.13) is a fundamental Wold representation of the equilibrium time series process.

3.2.4 Confronting the identification problem

Fundamentalness of the structural shocks with respect to the econometrician’s information set is necessary but not sufficient to be able to correctly estimate of their effects. In particular, after obtaining reduced-form estimates, the econometrician needs to recover the structural shocks through an appropriate identification strategy. The
econometrician thus estimates an unrestricted VAR of the form

\[ y_t^* = B_1 y_{t-1}^* + B_2 y_{t-2}^* + B_3 y_{t-3}^* + \cdots + u_t = C(L) u_t, \quad u_t \sim N(0, \Sigma), \]  

(3.14)

where the \( B_i, i = 1, 2, 3, \ldots \), are matrices of coefficients and \( C(L) \) is the infinite order multivariate lag polynomial of the moving average representation in the innovations \( u_t \), satisfying \( C(0) = I \). Assuming that there exists a linear mapping between the reduced-form innovations and the structural shocks, i.e. \( u_t = D \epsilon_t \), the moving average representation in the structural shocks is given by \( y_t^* = C(L) D \epsilon_t \) with \( \epsilon_t = D^{-1} u_t \).

Normalizing \( \text{cov} (\epsilon_t) = I \), the impact matrix \( D \) must satisfy \( DD' = \Sigma \) or equivalently \( D = AR \), where \( R \) is an orthonormal matrix such that \( RR' = I \) and \( A \) is an arbitrary orthogonalization of the reduced-form covariance matrix \( \Sigma \). The latter could for example be achieved by a Cholesky decomposition of \( \Sigma \) such that \( R = I \).

Suppose that the econometrician is only interested in the effects of government spending shocks. When observing \( \{ \hat{g}_{t-j}, E_{t-j}\hat{g}_{t-j+2-j}, \hat{k}_{t-j}\}_{j=0}^\infty \), the econometrician could identify the two types of spending shocks by taking the innovations to spending as the unanticipated shocks and by taking the innovations to expected spending that are orthogonal to the latter as the anticipated shocks. The shocks can thus be identified by a Cholesky decomposition of the reduced-form covariance matrix associated with (3.14) that has government spending ordered before its two-quarter ahead expectation and with capital ordered last. In fact, the impact matrix that is obtained by setting \( L = 0 \) in (3.13) is lower triangular:

\[
P^*(0) = \begin{bmatrix}
1 & 0 & 0 \\
\rho_g^2 & 1 & 0 \\
\eta_{kg} & \eta_{k\epsilon,1} & \eta_{ka}
\end{bmatrix},
\]

which implies that the processes (3.13) and (3.14) have a Cholesky structure.

As an alternative, the econometrician could observe the one-period expectational error \( \hat{g}_t - E_{t-1}\hat{g}_t \) instead of realized spending \( \hat{g}_t \). Since \( E_{t-1}\hat{g}_t = \rho_g \hat{g}_{t-1} + \epsilon_{t-2}^u \), it follows that \( \hat{g}_t - E_{t-1}\hat{g}_t = \epsilon_t^u \). The VAR model in \( \{ \hat{g}_{t-j} - E_{t-j}\hat{g}_{t-j}, E_{t-j}\hat{g}_{t-j+2-j}, \hat{k}_{t-j}\}_{j=0}^\infty \) is
then given by

\[
\begin{bmatrix}
\hat{g}_t - E_{t-1}\hat{g}_t \\
E_t\hat{g}_{t+2} \\
\hat{k}_t
\end{bmatrix}
= \begin{bmatrix}
1 & 0 & 0 \\
\frac{\rho_g^2}{1-\rho_g L} & \frac{1}{1-\rho_g L} & 0 \\
\frac{\eta_g}{(1-\eta_k L)(1-\rho_g L)} & \frac{\eta_g L^2+\eta_{kk,2}(1-\rho_g L)(\theta_L+L)}{(1-\eta_k L)(1-\rho_g L)} & \frac{\eta_k}{(1-\eta_k L)(1-\rho_a L)}
\end{bmatrix}
\begin{bmatrix}
\varepsilon_{u,g,t} \\
\varepsilon_{a,g,t} \\
\varepsilon_{a,t}
\end{bmatrix},
\]

or

\[
y^{**}_t = P^{**}(L)\varepsilon_t.
\]

The determinant of \( P^{**}(z) \) is given by

\[
\det P^{**}(z) = \frac{\eta_{ka}}{(1-\rho_g z)(1-\eta_{kk} z)(1-\rho_a z)},
\]

such that (3.15) is also an invertible moving average process. The econometrician could now identify the spending shocks by taking the expectational errors as the unanticipated shocks and by taking the innovations to expected spending that are orthogonal to the expectational errors as the anticipated shocks. Accordingly, the shocks can again be identified by a Cholesky decomposition as the impact matrix associated with (3.15) is also lower triangular; in fact, it satisfies \( P^{**}(0) = P^*(0) \) as given above.

The two identification strategies just discussed are however not equivalent: the first variant relies on VAR forecasts to achieve identification while the second variant uses additional information from data on expectational errors. The following section thus compares the two variants using stochastic simulations.

### 3.3 Simulation evidence

This section tests the usefulness of the proposed identification strategies. In particular, model-based stochastic simulations are conducted to compare the expectations-based approach and the standard recursive SVAR approach, which does not take into account the possibility of policy foresight. Section 3.4 discusses modifications to the benchmark model to check the robustness of the expectations-based approach, for example to alternative assumptions on the government spending process.
3.3.1 Monte Carlo set-up

The approach implemented here is a Monte Carlo exercise, following e.g. Ramey (2011b). That is, $M$ data samples of length $T$ are generated from the calibrated model.$^6$ The different approaches are first evaluated in terms of their large-sample properties, setting $T = 10,000$ and $M = 100$. Small-sample results are discussed in Section 3.4. The estimated impulse responses for each of the $M$ samples are ordered and the mean estimates are reported, with 90 percent two-sided error bands. The estimated responses are then compared to the impulse responses implied by the model.

3.3.2 Standard SVAR identification

The properties of the standard SVAR identification approach are investigated first. Government spending, TFP, and capital are thus taken as observable and included in this order in the VAR model. As shown above, the equilibrium time series process associated with these observables has a non-fundamental equilibrium representation. Investment is added as a fourth variable to the regression equation. As there are only three shocks, to avoid stochastic singularity a measurement error on investment $\varepsilon_{i,t} \sim N(0, \sigma_i^2)$ with $\sigma_i = 0.0001$ is included in the data-generating process (DGP).

Figure 3.4 shows the estimated impulse responses for the benchmark calibration, where a government spending shock is identified as the innovation to government spending using a Cholesky decomposition of the reduced-form covariance matrix. The figure also shows the impulse responses to both shocks implied by the DGP from the quarter in which spending increases onwards. The impulse responses of course cannot pick up any variation in investment and capital due to anticipated shocks during the anticipation period. However, the results seem to suggest that the error with respect to the effects of surprise spending shock is not very large. The possibility of policy foresight may thus not matter much quantitatively even if it is ignored.

However, the latter is not true in general. Figure 3.5 reports the estimated impulse responses when the anticipation rate $\theta$ is reduced to 0.58 by reducing the subjective

$^6$In this section, for convenience, the log-linearized model is solved numerically by the Gensys algorithm (see Sims, 2004).
Figure 3.4: Monte Carlo impulse responses – standard SVAR scheme I

Notes. Benchmark calibration ($\theta = 0.93$); SVAR responses (means) and 90 percent error bands are based on 100 samples of 10,000 observations each; the spending shock is identified by ordering government spending first in a Cholesky decomposition; DGP responses to anticipated shock are plotted from spending increase onwards; responses are measured as relative percentage deviations from steady state.

discount factor $\beta$ to 0.8.\textsuperscript{7} The results show that the effects on investment and capital of neither of the two shocks are correctly estimated. The SVAR responses indicate an initial decline in investment followed by an increase, whereas the actual investment response to the unanticipated shock is positive for several quarters while the response to the anticipated shocks is uniformly negative (after the anticipation period). There is also a relatively strong downward bias in the estimated spending response, such that

\textsuperscript{7}This is of course a relatively low value for $\beta$; alternatively, one could reduce the anticipation rate by increasing the intertemporal elasticity of substitution (which is equal to one with log utility) or lowering the capital share in production $\alpha$ (see Mertens and Ravn, 2010). A longer anticipation horizon would also create a stronger wedge between the effects of anticipated and unanticipated shocks.
the econometrician would overstate the overall expansionary effect of government expenditures on investment. Hence, there are realistic cases where the standard recursive SVAR identification approach can lead to misleading conclusions.\footnote{The bias in the estimated impulse responses turns even larger, and more in line with Ramey’s (2011b) findings on the effects of anticipated shocks, if the standard deviation of anticipated shocks is increased relative to the standard deviation of unanticipated shocks (not reported).}

### 3.3.3 Expectations-based approach

The properties of the expectations-based identification approach are discussed next. The variant with expectational errors is analyzed first. The econometrician thus estimates a VAR in \( \{ \hat{g}_{t-j} - \hat{g}_{t-j}, E_{t-j} \hat{g}_{t+2-j}, \hat{k}_{t-j}, \hat{i}_{t-j}^{obs} \}_j \). Adding investment with a measurement error to the regression goes without prejudice to the results discussed.
Notes. Benchmark calibration ($\theta = 0.93$); an unanticipated spending shock is identified by ordering the expectational error on spending first in a Cholesky decomposition, the shock is a one percent increase in the expectational error in quarter 0; an anticipated spending shock is identified by ordering the two-quarter ahead expectation of spending second, the shock being a one percent increase in the two-quarter ahead expectation in quarter 0; responses are measured as relative percentage deviations from steady state.

in the previous section. To see this, notice that the solution for investment is given by

$$\hat{\dot{i}}_t = \eta_{ik} \hat{k}_{t-1} + \eta_{ia} \hat{a}_t + \eta_{ig} \hat{g}_t + \eta_{i\varepsilon} \varepsilon_{g,t-1}^a + \eta_{i\varepsilon} \varepsilon_{g,t}^a,$$

and observed investment is $\dot{i}^{obs}_t = \hat{\dot{i}}_t + \varepsilon_{i,t}$. Hence, the econometrician’s VAR reads

$$\begin{bmatrix} \dot{g}_t - E_{t-1} \hat{g}_t \\ E_t g_{t+2} \\ \hat{k}_t \\ \dot{i}^{obs}_t \end{bmatrix} = \begin{bmatrix} 1 \\ \rho_s^2 \\ \frac{\eta_{ik}}{1 - \rho_g L} \\ \frac{\eta_{ka} L^2 + \eta_{i\varepsilon} \Theta_2(L)}{1 - \rho_a L} \\ \frac{-\eta_{ig}}{1 - \rho_a L} \\ \Theta_1(L) \\ \Theta_2(L) \\ \Theta_3(L) \end{bmatrix} \begin{bmatrix} \varepsilon_{g,t}^u \\ \varepsilon_{g,t}^a \\ \varepsilon_{a,t} \\ \varepsilon_{i,t} \end{bmatrix}.$$
Figure 3.7: Monte Carlo impulse responses – expectations-based scheme II

Notes. Lower anticipation rate \((\theta = 0.58, \beta = 0.8)\); see Figure 3.6.

where the \(\Theta_s(L), s = 1, 2, 3\), follow from substituting out capital, TFP, and spending in (3.16). The determinant of the lag matrix is again equal to \(\eta_{ka}/[(1 - \rho_g z)(1 - \eta_{kk} z)(1 - \rho_a z)]\) as for (3.15), such that this modified process is also a fundamental one.

Figure 3.6 reports the expectations-based SVAR impulse responses to both types of spending shocks. The results show that the estimated effects closely match those of the DGP. Under the unanticipated shock, investment declines and the impact response of the two-quarter ahead expectation (expected spending in quarter \(t + 2\) conditional on quarter \(t\) information) is close to the theoretical value of \(\rho_g^2 = 0.72\). Under the anticipated shock, the two-quarter ahead expectation of spending increases by one percent in quarter 0. Importantly, investment increases during the anticipation period but it is negative from quarter 2 onwards, as implied by the DGP.

The previous exercise is now repeated when the anticipation rate \(\theta\) is reduced
Notes. Lower anticipation rate ($\theta = 0.58$, $\beta = 0.8$); an unanticipated spending shock is identified by ordering spending first in a Cholesky decomposition, the shock is a one percent increase in spending in quarter 0; an anticipated spending shock is identified by ordering the two-quarter ahead expectation of spending second, the shock being a one percent increase in the two-quarter ahead expectation in quarter 0; responses are measured as relative percentage deviations from steady state.

to 0.58 by reducing $\beta$ to 0.8, which was seen to increase the relevance of the non-fundamentalness problem of the standard recursive SVAR approach. However, Figure 3.7 shows that for the expectations-based approach the results are robust to changes in $\theta$: the estimated effects of both shocks are similarly close to the DGP effects as under the benchmark calibration. The identification approach is also robust to changes in the relative volatility of the two types of spending shocks (not reported).

The variant of the expectations-based approach where actual spending is observed instead of the expectational errors is analyzed next. For brevity, only results for the lower anticipation rate are reported in Figure 3.8. The figure reveals some noticeable
differences in the SVAR impulse responses and the DGP responses; in particular, the estimated investment response to the unanticipated shock is not as close to DGP response as under the variant with expectational errors. The additional information from data on expectational errors therefore seems useful to obtain precise estimates. The analysis thus proceeds with that variant, but Section 3.5 also checks the robustness of the empirical results when the variant with spending is used to achieve identification.

3.4 Robustness

This section discusses the results of three types of robustness exercises. First, the experiment of the previous section is repeated for a smaller sample size. Second, the spending process is modified by allowing feedbacks from other economic shocks (TFP shocks in the present model) on spending. Third, it is checked whether surprise spending shocks can also be correctly identified in a VAR model that does not include expectations on future spending but only expectational errors, as in Ramey (2011b) and Auerbach and Gorodnichenko (2010). The implications for the empirical application are discussed at the end of this section.

3.4.1 Small sample results

The results reported so far were based on large samples ($T = 10,000$). For Figure 3.9, the Monte Carlo exercise is repeated for an empirically realistic sample size of $T = 114$ and $M = 10,000$. The reduction in the sample size implies that the data contains less information, so the error bands become wider. However, the point estimates remain close to the DGP responses. The bias of the standard SVAR approach is therefore still eliminated by the expectations-based approach.

3.4.2 Spending reaction to lagged TFP

Consider now the following modification of the spending process (3.4):

$$\log(g_t/\bar{g}) = \rho_g \log(g_{t-1}/\bar{g}) + \rho_{ga} \log a_{t-1} + \varepsilon_{g,t}^a + \varepsilon_{g,t-2}, \quad \rho_{ga} \in \mathbb{R}. \tag{3.17}$$

The sample size is equal to the data sample below, i.e. 114 quarters from 1981Q4 to 2010Q1.
Figure 3.9: Robustness I – Monte Carlo impulse responses in small samples

Notes. Sample size is \( T = 114 \); see Figure 3.6.

According to (3.17), government spending reacts with a one-period lag to fluctuations in productivity. This modification is a convenient shortcut for more complicated endogenous feedbacks on government spending (e.g. from movements in output or government revenues) that allows to obtain simple analytical expressions. Impulse responses to one percent productivity shocks in the model with (3.17) replacing (3.4) are shown in Figure 3.10. Absent any spending feedback (\( \rho_{ga} = 0 \), left panel), the shock increases consumption, hours, output, and investment. Hours increase because the substitution effect from higher productivity is larger than the positive wealth effect of higher productivity on lifetime income. When spending increases with TFP (\( \rho_{ga} = 1 \), right panel), there is a smaller wealth effect due to government absorption of goods and services, so the increase in hours and output (consumption) is stronger (weaker).

Under the modified spending process, the one-period expectational error is still
Figure 3.10: Robustness II – model impulse responses to a productivity shock when spending can react to lagged productivity

Notes. Both panels show responses to one percent surprise increases in TFP relative to its steady state value; left panel: no spending reaction to TFP ($\rho_{ga} = 0$); right panel: procyclical spending reaction to lagged TFP ($\rho_{ga} = 1$); responses are percentage deviations from steady state.

given by $E_{t-1}\hat{g}_t - \hat{g}_t = \varepsilon_{g,t}^u$, since spending only reacts with a lag to productivity shocks. However, two-quarter ahead expected spending becomes

$$E_t\hat{g}_{t+2} = \frac{\rho_g^2}{1 - \rho_g L} \varepsilon_{g,t}^u + \frac{1}{1 - \rho_g L} \varepsilon_{g,t}^a + \frac{\rho_{ga} \rho_a}{1 - \rho_a L} \varepsilon_{a,t},$$

such that expected spending is affected by the current state of productivity. Suppose that the econometrician estimates a similar VAR as above:

$$\begin{bmatrix}
\hat{g}_t - E_{t-1}\hat{g}_t \\
E_t\hat{g}_{t+2} \\
s_{obs,t} \\
\hat{g}_t
\end{bmatrix} =
\begin{bmatrix}
1 & 0 & 0 & 0 \\
\frac{\rho_g^2}{1 - \rho_g L} \eta_{ka} & \frac{1}{1 - \rho_g L} \eta_{ka} \eta_{az}(1 - \rho_a L) & \frac{\rho_{ga} \rho_a}{1 - \rho_a L} & \frac{\rho_{ga} \rho_a}{1 - \rho_a L} \\
\frac{1}{(1 - \eta_{ka}\L)(1 - \rho_g L)} \Theta_1(L) & \frac{1}{(1 - \eta_{ka}\L)(1 - \rho_g L)} \Theta_2(L) & 0 & \frac{\rho_{ga} \rho_a}{1 - \rho_a L} \\
\frac{1}{(1 - \eta_{ka}\L)(1 - \rho_g L)} & \frac{1}{(1 - \eta_{ka}\L)(1 - \rho_g L)} & 0 & \Theta_3(L) \\
0 & 0 & 0 & 0
\end{bmatrix}
\begin{bmatrix}
\varepsilon_{g,t}^u \\
\varepsilon_{g,t}^a \\
\varepsilon_{a,t} \\
\varepsilon_{i,t}
\end{bmatrix}.$$  

The determinant of the lag matrix is equal to $\eta_{ka}^*/[(1 - \rho_g z)(1 - \eta_{ka}z)(1 - \rho_a z)]$, such that the process is fundamental.\(^{(10)}\) However, the presence of the term $\rho_{ga} \rho_a/(1 - \rho_a L)$ makes the identification problem more difficult: the econometrician cannot distinguish changes in expected spending due to anticipated spending shocks and TFP shocks by

\(^{(10)}\)Notice the presence of $\eta_{ka}^*$, which is equal to $\eta_{ka}$ only for $\rho_{ga} = 0$. The coefficient $\eta_{ca}$ also changes to $\eta_{ca}^*$, and similarly for $\Theta_3(L)$ which becomes $\Theta_3^*(L)$. Details are provided in Appendix 3.A.
Figure 3.11: Robustness III – Monte Carlo impulse responses to an anticipated spending shock when spending reacts to lagged productivity

Figure 3.11 demonstrates the implications of the missing Cholesky structure, if the econometrician nevertheless attempts to estimate the effects of anticipated spending shocks by identifying the latter as increases in expected spending that are orthogonal to expectational errors. The estimated effects are seen to be located in between the responses to anticipated spending shocks and TFP shocks implied by the DGP. Of course, the bias becomes smaller with a smaller reaction of spending to the state of productivity. However, the identification scheme produces a bias even for relatively small feedbacks $\rho_{ga}$; for negative $\rho_{ga}$, the bias turns negative.

One way to address those issues is to condition on TFP in the VAR model. That

Notes. Spending reacts procyclically to lagged TFP ($\rho_{ga} = 1$); see Figure 3.6.
Figure 3.12: Robustness IV – Monte Carlo impulse responses to an anticipated spending shock when spending reacts to lagged productivity (observed)

Notes. Spending reacts procyclically to lagged TFP ($\rho_{ga} = 1$); an anticipated spending shock is identified by ordering the two-quarter ahead expectation of spending third in a Cholesky decomposition, the shock being a one percent increase in the two-quarter ahead expectation in quarter 0; the expectational error on spending is ordered second and TFP is ordered first; responses are measured as relative percentage deviations from steady state.

is, suppose that the econometrician includes $\hat{a}_t$ as the first variable in the VAR:

$$
\begin{bmatrix}
\hat{a}_t \\
\hat{g}_t - E_{t-1}\hat{g}_t \\
E_{t+1}\hat{g}_{t+2} \\
\hat{k}_{t,obs}
\end{bmatrix} =
\begin{bmatrix}
\frac{1}{1-\rho_a L} & 0 & 0 & 0 \\
0 & 1 & 0 & 0 \\
\frac{\rho_{ga}\rho_g}{(1-\rho_a L)(1-\rho_g L)} & \frac{\rho_g^2}{1-\rho_g L} & \frac{\eta_{ka}}{(1-\rho_a L)(1-\rho_g L)} & \frac{1}{1-\rho_g L} \\
\eta_a \eta_{ka} & \frac{\eta_a}{(1-\rho_a L)(1-\rho_g L)} & \frac{\eta_{ka} L^2 + \eta_{ka} \rho_g (1-\rho_a L)(\theta + L)}{(1-\rho_a L)(1-\rho_g L)} & 1
\end{bmatrix}
\begin{bmatrix}
\varepsilon_{a,t} \\
\varepsilon_{g,t} \\
\varepsilon_{g,t} \\
\varepsilon_{k,t}
\end{bmatrix}.
$$

Notice that investment and its measurement error have been dropped and instead there is a measurement error on capital, $\varepsilon_{k,t} \sim N(0, \sigma_k^2)$ with $\sigma_k = 0.0001$. Furthermore, the surprise spending shock is now ordered second and the news shock is ordered third.
This is again a fundamental process, the determinant of the lag matrix being equal to 
\[(1 - \rho_L)(1 - \rho_a L)^{-1},\] and the impact matrix has a Cholesky structure.

Figure 3.12 shows that, by conditioning on TFP, the anticipated spending shock is 
again well identified. However, the requirements on the econometrician’s information 
set have become more stringent under this modified identification scheme, since TFP 
needs to be available as an observable variable for the scheme to work.

### 3.4.3 Spending reaction to current TFP

If there is a contemporaneous feedback from TFP on spending, the expectational error 
on spending is a weighted average of unanticipated spending shocks and TFP shocks, 
where the weight on TFP shocks is given by the strength of the feedback. That is, if

\[
\log(g_t/\bar{g}) = \rho_g \log(g_{t-1}/\bar{g}) + \rho_{ga} \log a_t + \varepsilon_{u,t}^g + \varepsilon_{a,t-2}^a, \quad \rho_{ga} \in \mathbb{R}, \tag{3.18}
\]

the expectational error on spending is a mix of TFP shocks and surprise spending 
shocks: \(\hat{y}_t - E_{t-1}\hat{g}_t = \rho_{ga}\varepsilon_{a,t} + \varepsilon_{u,t}^g\). If TFP remains unobserved, the identification of 
surprise spending shocks based on expectational errors would therefore fail. This is 
demonstrated in Figure 3.13, which shows the estimated responses to an unanticipated 
spending shock that is identified by the associated expectations-based scheme. Similarly as above, the estimated responses are biased as they are located in between the 
responses to spending shocks and TFP shocks implied by the DGP.

However, if TFP can be observed, the econometrician could estimate the model

\[
\begin{bmatrix}
\hat{a}_t \\
\hat{g}_t - E_{t-1}\hat{g}_t \\
E_t\hat{g}_{t+2} \\
k_t^{obs}
\end{bmatrix} = \begin{bmatrix}
\frac{1}{1-\rho_a L} & 0 & 0 & 0 \\
\rho_g a_t & 1 & 0 & 0 \\
\frac{\rho_{ga}^2}{1-\rho_a^2 L} & \frac{\rho_g^2}{1-\rho_a L} & 1 & 0 \\
\frac{\eta_{ga}}{(1-\eta_{ga})(1-\rho_a L)} & \frac{\eta_{ga}}{(1-\eta_{ga})(1-\rho_a L)} & \frac{n_{k,t}^2+\eta_{ka} L(1-\rho_a L)(\theta+L)}{(1-\eta_{ka})(1-\rho_a L)} & 1
\end{bmatrix} \begin{bmatrix}
\varepsilon_{a,t} \\
\varepsilon_{u,t}^g \\
\varepsilon_{a,t-2}^a \\
\varepsilon_{k,t}
\end{bmatrix},
\]

and apply the expectations-based identification scheme, conditioning on TFP when 
estimating the effects of spending shocks. The estimated impulse responses to a surprise 
spending shock are shown in Figure 3.14. The results show that the effects of the shock 
are well identified under the adjusted scheme.
Figure 3.13: Robustness V – Monte Carlo impulse responses to an unanticipated spending shock when spending reacts to current productivity

Notes. Spending reacts procyclically to current TFP ($\rho_{ga} = 1$); see Figure 3.6.

3.4.4 Surprise shocks under non-fundamentalness

Ramey (2011b) and Auerbach and Gorodnichenko (2010) suggest to extract the unanticipated component of exogenous movements in government spending through expectational errors on spending. In both of these studies an innovation to the forecast error is interpreted as an unanticipated spending shock. However, the studies do not include expectations on future spending in the regression. It can be shown, similarly as above, that a VAR that is specified in this way is not fundamental. Would an econometrician who applies this type of “non-fundamental” identification strategy still correctly estimate the effects of unanticipated spending shocks?

Figure 3.15 compares point estimates for ten simulated data sets from the bench-
Figure 3.14: Robustness VI – Monte Carlo impulse responses to an unanticipated spending shock when spending reacts to current productivity (observed)

Notes. Spending reacts procyclically to current TFP ($\rho_{ga} = 1$); an unanticipated spending shock is identified by ordering the expectational error on spending second in a Cholesky decomposition, the shock being a one percent increase in the expectational error in quarter 0; the two-quarter ahead expectation of spending is ordered third and TFP is ordered first; responses are measured as relative percentage deviations from steady state.

mark specification of the model that was used for Figure 3.6. The left-hand panels show the estimates under the expectations-based scheme, where the two-quarter ahead expectation of spending is included in the VAR. The right-hand panels show the estimates under the non-fundamental scheme, where expected spending is not included in the regression. The estimated regressions include otherwise identical simulated data. In both cases, consumption is added as an additional observed variable with a small measurement error of 0.001 on observed consumption in the model. Unlike the fully expectations-based identification, the non-fundamental identification produces a downward bias in the estimated responses of investment, capital, and consumption,
Figure 3.15: Robustness VII – Monte Carlo impulse responses under non-fundamental scheme

Notes. Benchmark calibration; an unanticipated spending shock is identified by ordering the expectational error on spending first in a Cholesky decomposition, the shock being a one percent increase in the expectational error in quarter 0; left panels: two-quarter ahead expectation of spending is included as the second variable; right panels: two-quarter ahead expectation is not included; responses are measured as relative percentage deviations from steady state.

especially at longer horizons. The results of this exercise therefore suggest that an SVAR identification strategy based on expectational errors only is not appropriate to estimate the effects of surprise spending shocks.

3.4.5 Implications for applied research

What are the implications of the above findings for applied research on the macroeconomic effects of government spending?
First, the results show that an expectations-based identification approach can help to solve the non-fundamentalness problem that distorts econometric inference under the standard recursive SVAR identification approach. For standard information flows, the sufficient condition is that expectations on future spending up the anticipation horizon of economic agents should be included in the VAR.

Second, with respect to the identification problem of distinguishing anticipated spending shocks from unanticipated spending shocks and other economic shocks, the conclusions are mixed. If spending is affected by other shocks, these shocks need to be known, observed, and conditioned upon. However, there is significant uncertainty on which shocks do affect government spending. In addition, most structural shocks are unobserved state variables, so they cannot be included in the econometrician’s information set.11 The expectations-based identification of both types of spending shocks is therefore prone to significant problems.12

The good news is that, by exploiting variation in expectational errors, surprise spending shocks can be robustly identified if expected future spending is included in the VAR and if spending only reacts with some lag to other economic shocks. Since government spending is usually defined as government final consumption plus government investment and thus net of transfer and interest payments, the assumption that spending does not react within a quarter to other shocks seems justified. Government spending is then arguably acyclical, such that there is no automatic reaction of spending to movements in the business cycle, and a discretionary fiscal response to economic shocks is unlikely to occur within a quarter due to implementation lags in the policy process. Overall, the expectations-based approach is thus found to be useful for the identification of surprise spending shocks when it includes standard short-run exogeneity restrictions.

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11 Few exceptions such as multifactor productivity estimates are available from official sources but usually only on an annual basis (e.g. U.S. productivity estimates from the Bureau of Labor Statistics).
12 Of course, things are even worse for the econometrician if there is not only news on future fiscal variables but also on future economic shocks (e.g. productivity news, see Beaudry and Portier, 2006) to which spending might react in the future. The econometrician would then need to condition on expectations of future unobserved state variables!
3.5 Empirical application

This section discusses the results of an empirical application to the U.S. that uses survey data on federal government spending from the Survey of Professional Forecasters to measure economic agents’ (or market participants’) expectations. Given the findings of the previous sections, the empirical application focuses on the effects of surprise spending shocks. The discussion further concentrates on a comparison of estimates from the expectations-based approach and the standard recursive SVAR identification approach (see e.g. Blanchard and Perotti, 2002; Fatás and Mihov, 2001; Perotti, 2005), which does not take into account the possibility of policy foresight.

3.5.1 Data description

Figure 3.16 shows the deviations of real federal government spending from the predictions of the respondents to the Philadelphia Fed’s Survey of Professional Forecasters over the period 1981Q4 to 2010Q1. Government spending is defined as the sum of government consumption and gross investment. Details on data definitions are provided in Appendix 3.B. The expectational errors are computed on the basis of the average predictions across all panelists made one and four quarters earlier. The forecasts submitted in quarter \( t \) are also taken conditional on quarter \( t \) information, although the official documentation of the survey takes forecasts made in \( t \) conditional on \( t - 1 \) information.\(^\text{13}\) The reason for the latter is that the questionnaires are sent out right after the advance report of the Bureau of Economic Analysis (BEA) is released, which contains the first estimates of GDP and its components for the previous quarter. However, the forecasters form their expectations conditional on all information which they have available in period \( t \), and which is not necessarily restricted to the BEA report. Conditioning on the information set at the time when the forecast is made thus seems reasonable.\(^\text{14}\)

\(^{13}\)The official documentation is available from http://www.philadelphiahfed.org.

\(^{14}\)The forecasts for levels were originally scaled to the national accounts base year that had been in effect at the time the survey questionnaire was sent to the forecasters. Over time, as benchmark revisions to the data occur, the scale changes. As there have been a number of base year changes in U.S. national accounts since the survey began, the forecasts were therefore scaled to the current base year, 2005, through backcasting by the actual growth rates and imposing the average growth rate over the sample at the break points.
Figure 3.16: Expectational errors on U.S. federal government spending

Notes. Quarterly data, period 1981Q4-2010Q1; data on expectations is taken from Philadelphia Fed’s Survey of Professional Forecasters; expectational errors are computed as log differences (in percent) of real spending in quarter $t$ and the prediction of real spending made in quarters $t - 1$ (thick line) and $t - 4$ (thin line).

3.5.2 Expectations-based identification

The expectational errors shown in Figure 3.16 indicate the presence of a pronounced unanticipated component in federal government spending. A natural next step is to exploit this variation to estimate the effects of surprise spending shocks by an application of the identification strategy analyzed above. To achieve fundamentalness, expectations on future government spending are included in the VARs in log-levels. As the precise anticipation horizon of economic agents is uncertain, two reduced-form VARs are estimated by OLS which include, respectively, the two- and four-quarter ahead expectations of spending. Real GDP, private consumption, and private investment are added
Figure 3.17: Empirical impulse responses I – expectations-based scheme

Notes. Normalized one percent increase in expectational error on federal government spending in quarter 0; the VAR models include the one-quarter expectational errors as the first variable and the two-quarter (left panels) and four-quarter (right panels) ahead expectations of spending as the second variable (responses not shown); 90 percent two-sided confidence bands around mean impulse responses are reported, calculated by 1,000 bootstrap replications; data definitions are provided in Appendix 3.B.

as additional variables in log-levels. Both VARs include four lags of the endogenous variables, a constant, and a quadratic time trend.\textsuperscript{15}

Surprise spending shocks are then identified as the innovations to expectational errors that have a contemporaneous impact on all other variables in a system with expected spending, output, consumption, and investment, by a Cholesky decomposition of the estimated reduced-form covariance matrix. Figure 3.17 shows the estimated mean impulse responses of the expectational errors, output, consumption, and investment to one percent shocks of this type. The figure also shows 90 percent two-sided

\textsuperscript{15}The results are robust to the use of a linear trend and three or five lags (not reported).
Notes. Normalized one percent increase in federal government spending in quarter 0; the VAR model does not include expectations data; the spending shock is identified by ordering spending before the remaining variables in a Cholesky decomposition; 90 percent two-sided confidence bands around mean impulse responses are reported, calculated by 1,000 bootstrap replications; data definitions are provided in Appendix 3.B.

bootstrap confidence bands for the estimated impulse responses. According to both VARs, a surprise spending shock leads, on average, to an initial increase in output. Consumption and investment hardly react on impact but start to decline shortly after the shock. After a few quarters, both consumption and investment turn significantly negative, which is associated with a reversal of the output effect.

3.5.3 Comparison with standard SVAR identification

Figure 3.18 reports the results of an application of the standard recursive SVAR identification scheme, according to which government spending shocks are identified by a
Figure 3.19: Empirical impulse responses III – both schemes

Notes. The expectations-based VAR models (solid lines) include the one-quarter expectational errors (ordered first), realized spending (ordered second), and the two-/four-quarter ahead expectations of spending (ordered third); surprise spending shocks are identified as in Figure 3.17 and the standard SVAR scheme (dashed lines) goes as in Figure 3.18; mean estimated impulse responses are reported.

Cholesky decomposition as the innovations to spending that have a contemporaneous impact on output, consumption, and investment. In contrast to the previous results, a shock that is identified in this way leads to increases in consumption and investment during two quarters, which are however not significant at the 90 percent level. In addition, both impulse responses and also the output response are more persistent than under the expectations-based approach; in particular, they do not turn significantly negative until towards the end of the horizon considered.

The expectations-based VAR models estimated for Figure 3.17 do not include the level of government spending as an endogenous variable but the standard VAR model
Table 3.2: Government spending multipliers\textsuperscript{a}

<table>
<thead>
<tr>
<th></th>
<th>Impact</th>
<th>4 qrts.</th>
<th>8 qrts.</th>
<th>12 qrts.</th>
<th>20 qrts.</th>
</tr>
</thead>
<tbody>
<tr>
<td>\textit{Two quarters anticipation}\textsuperscript{b}</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>GDP</td>
<td>1.10</td>
<td>-1.57</td>
<td>-1.02</td>
<td>-1.28</td>
<td>-1.29</td>
</tr>
<tr>
<td>Spending</td>
<td>1.00</td>
<td>0.59</td>
<td>0.36</td>
<td>0.24</td>
<td>-0.02</td>
</tr>
<tr>
<td>\textit{Four quarters anticipation}\textsuperscript{b}</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>GDP</td>
<td>1.07</td>
<td>-1.37</td>
<td>-0.79</td>
<td>-1.09</td>
<td>-1.44</td>
</tr>
<tr>
<td>Spending</td>
<td>1.00</td>
<td>0.64</td>
<td>0.40</td>
<td>0.24</td>
<td>0.00</td>
</tr>
<tr>
<td>\textit{Standard SVAR identification}</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>GDP</td>
<td>1.06</td>
<td>-0.57</td>
<td>-0.37</td>
<td>-0.66</td>
<td>-0.95</td>
</tr>
<tr>
<td>Spending</td>
<td>1.00</td>
<td>0.52</td>
<td>0.40</td>
<td>0.27</td>
<td>0.07</td>
</tr>
</tbody>
</table>

\textsuperscript{a} The multipliers on GDP are computed based on mean impulse responses according to the following formula: multiplier in quarter \( t = \text{GDP response in quarter } t/\text{(spending response in quarter 0 times average share of spending over GDP over the sample)}.\)

\textsuperscript{b} For the expectations-based identification, the multipliers are computed from two different VAR models which include the two-quarter and four-quarter ahead expectations of government spending, respectively.

does. Thus, to make the results comparable, two additional regressions are estimated where the level of spending is added (ordered second) next to the expectational errors and the expectations of future spending. The point estimates of the impulse responses from those models are compared to the point estimates from the standard VAR model in Figure 3.19. The results show that, despite similar spending responses in terms of size and persistence, the effects on output, consumption, and investment are, on average, substantially smaller under the expectations-based identification scheme.

To compare the magnitudes of the estimated fiscal multipliers, following Blanchard and Perotti (2002), Table 3.2 compares the dollar change in GDP due to the initial dollar change in government spending at different horizons for the VAR models from Figure 3.19. The entries in the rows for GDP can also be interpreted as multipliers on output due to a fiscal shock leading to an initial increase in the level of government spending of size 1% of GDP. The results show that both the expectations-based approach and the standard recursive SVAR approach yield multipliers on GDP of approximately 1.1 on impact. However, the expectations-based approach yields multipliers smaller than minus one at longer horizons, whereas the multipliers implied by the standard recursive
SVAR approach are uniformly larger than that.

### 3.5.4 Predictability of shocks

A possible explanation for the differences between the results from the two alternative approaches is that the impulse responses to the standard SVAR shocks may incorporate some of the effects of anticipated spending shocks. The standard recursive SVAR approach may then pick up the upward-sloping paths of the responses of consumption and investment to shocks that were anticipated some quarters in advance, whereas the econometrician treats the spending increase as if it was unanticipated. In fact, Ramey (2011b) shows that standard SVAR shocks for federal government spending are Granger-caused by professional forecasts made one to four quarters earlier; that is, the SVAR shocks are predictable. A likely implication of this finding is that by ignoring policy foresight the econometrician would not capture the true economic impact of discretionary changes in government spending, even if surprise spending shocks are the only object of interest.

Figure 3.20 compares the identified shocks from the two approaches for the period 2001Q1 to 2010Q1. During this period, three easily recognizable events are likely to have affected U.S. federal government spending. The first two are the wars in Afghanistan and Iraq which began, respectively, on October 7, 2001 and March 20, 2003. The third event is the American Recovery and Reinvestment Act (ARRA) which was signed into law by President Obama on February 17, 2009. These events are marked by vertical lines in Figure 3.20. The figure shows that spending shocks are identified by all approaches immediately after the three events. However, the standard SVAR approach identifies sizeable positive shocks during about two years after the beginning of the war in Iraq, whereas the expectations-based approach does not identify any large surprise spending shocks during this period. Based on those results, one might suspect that some of the shocks identified by the standard SVAR approach were indeed anticipated by economic agents.

To investigate whether this is the case, following Ramey (2011b), it is analyzed whether the professional forecasts Granger-cause the identified shocks from the two approaches. In particular, a series of F-tests are performed where the unrestricted test...
Figure 3.20: Identified spending shocks

Notes. Period 2001Q1-2010Q1; the expectations-based VAR model includes two-quarter ahead expectations of spending; see Figures 3.17 and 3.18.

The equation takes the form

\[ x_t = a_0 + a_1 x_{t-1} + a_2 x_{t-2} + \cdots + a_p x_{t-p} + b_1 f_{t|t-1} + b_2 f_{t|t-2} + \cdots + b_h f_{t|t-h}, \]

and where the restricted test equation is given by

\[ x_t = a_0 + a_1 x_{t-1} + a_2 x_{t-2} + \cdots + a_p x_{t-p}, \]

where \( x_t \) denotes the identified shocks in quarter \( t \) and \( f_{t|t-1}, \ldots, f_{t|t-h} \) the log first differences of forecasts on real federal government spending made up to \( h \) quarters earlier. The null hypothesis is thus that the forecasts \( f_{t|t-1}, \ldots, f_{t|t-h} \) do not Granger-cause the shocks; that is, the null hypothesis states that the shocks are not actually
Table 3.3: Granger causality tests on identified shocks

<table>
<thead>
<tr>
<th>Independent variable(^{a,b})</th>
<th>F-statistic</th>
<th>5% critical value</th>
<th>p-value</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td></td>
<td>(p = 1)</td>
<td>(p = 4)</td>
</tr>
<tr>
<td><strong>Standard SVAR identification</strong></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>One-quarter ahead forecasts</td>
<td>1.87</td>
<td>3.36(^*)</td>
<td>3.93</td>
</tr>
<tr>
<td>Two-quarter ahead forecasts</td>
<td>2.02</td>
<td>2.21</td>
<td>3.93</td>
</tr>
<tr>
<td>Three-quarter ahead forecasts</td>
<td>5.52(^**)</td>
<td>5.77(^**)</td>
<td>3.93</td>
</tr>
<tr>
<td>Four-quarter ahead forecasts</td>
<td>3.50(^*)</td>
<td>3.07(^*)</td>
<td>3.93</td>
</tr>
<tr>
<td>All forecasts simultaneously</td>
<td>4.60(^**)</td>
<td>4.39(^**)</td>
<td>3.93</td>
</tr>
<tr>
<td><strong>Two quarters anticipation(^c)</strong></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Two-quarter ahead forecasts</td>
<td>0.00</td>
<td>0.10</td>
<td>3.93</td>
</tr>
<tr>
<td>All forecasts simultaneously</td>
<td>0.04</td>
<td>0.35</td>
<td>3.93</td>
</tr>
<tr>
<td><strong>Four quarters anticipation(^c)</strong></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Four-quarter ahead forecasts</td>
<td>0.96</td>
<td>1.19</td>
<td>3.93</td>
</tr>
<tr>
<td>All forecasts simultaneously</td>
<td>0.80</td>
<td>1.24</td>
<td>3.93</td>
</tr>
</tbody>
</table>

\(^a\) The dependent variables are the identified shocks in quarter \(t\), which are regressed on a constant, \(p\) own lags, and the log difference of forecasted spending for quarter \(t\) made one to four quarters earlier.

\(^b\) The null hypothesis is that the forecasts do not Granger-cause the shocks. \(^**\) indicates rejection of the null hypothesis at the five percent significance level, \(^*\) at the ten percent significance level.

\(^c\) For the expectations-based identification, the identified shocks from the VAR models which include the two-quarter and four-quarter ahead expectations of government spending are taken as dependent variables, respectively.

forecastable by the professional forecasters’ predictions.

Table 3.3 reports the test results when the forecasts are included individually and jointly as independent variables in the unrestricted regression, for both \(p = 1\) and \(p = 4\) lags of the dependent variables. The results show that, indeed, the standard SVAR shocks are on average forecastable by the professional forecasters’ predictions. When the forecasts are included jointly in the unrestricted regression, the null hypothesis that the forecasts do not Granger-cause the shocks is clearly rejected at the five percent significance level. On the other hand, the null hypothesis cannot be rejected for the shocks identified on the basis of expectational errors. These results suggest that the expectations-based approach is successful in extracting the unpredictable part of exogenous spending changes. Any bias of the standard SVAR approach should therefore be eliminated by the expectations-based identification scheme.
3.5.5 Extended regression specification

As a next step, the impact of surprise spending shocks on a broader set of indicators is investigated. That is, the real wage, the 3-month treasury bill rate, and the federal government debt-to-GDP ratio are added as additional endogenous variables in the regression. The real wage is added since it is an important variable in the controversy on the effects of government spending shocks.\textsuperscript{16} The T-bill rate is added to assess the impact of spending shocks on interest rates. The debt-to-GDP ratio is included to capture financing aspects and also to address the problem of omitted state variables of\textsuperscript{16}See, for instance, Linnemann and Schabert (2003), Perotti (2008), or Ramey (2011b).

\textit{Notes.} The VAR model includes two-quarter ahead expectations of spending; see Figure 3.17.
The results are reported in Figure 3.21. The estimated VAR model includes the two-quarter ahead expectation of spending and the level of spending.\footnote{The results are robust to using the four-quarter ahead expectation of spending instead.} Also according to the extended specification, the spending increase has small effects on output and leads to a decline in consumption and investment. The response of the real wage is insignificant in the short run and negative at longer horizons. The interest rate declines immediately. In terms of financing, the spending shock is associated with a sustained increase in the debt-to-GDP ratio. Hence, also in the extended specification a surprise spending shock is estimated to have contractionary effects over the medium term, in contrast to most of the previous SVAR literature but in line with the findings of, for instance, Mountford and Uhlig (2009) and Ramey (2011b).

### 3.5.6 Alternative identification scheme

As a final step of the analysis, the variant of the expectations-based scheme that relies on VAR forecasts in the identification of surprise spending shocks is applied. As discussed in Section 3.2.4, this scheme has government spending ordered first before expected spending as the second variable in the VAR model, and takes the innovations to spending that have a contemporaneous impact on all variables (by a Cholesky decomposition) as the unanticipated spending shocks.

The results of the application of this scheme in comparison to the standard recursive scheme are shown in Figure 3.22. The figure shows that the previous conclusions remain intact from a qualitative point of view: the estimated effects on output, consumption, and investment are, on average, substantially smaller under the expectations-based scheme than under the standard recursive scheme. Quantitatively, the effects are even weaker than under the benchmark expectations-based scheme especially at longer horizons (cf. Figure 3.19). The latter strengthens the conclusion that unanticipated discretionary changes in federal government spending were not very effective in raising U.S. economic activity over the period considered.
Figure 3.22: Empirical impulse responses V – alternative scheme

Notes. The expectations-based VAR models (solid lines) include realized spending (ordered first), and the two-/four-quarter ahead expectations of spending (ordered second); the alternative scheme identifies surprise spending shocks as one percent increases in spending by a Cholesky decomposition in this model; see Figure 3.18 for the standard SVAR scheme (dashed lines); mean estimated impulse responses are reported.

3.6 Conclusion

This chapter has demonstrated how the econometric problems created by foresight on government spending can be addressed when economic agents’ expectations on future spending can be incorporated, e.g. through survey data, in a VAR model. By a combination of theory and stochastic simulations, the chapter has shown that incorporating expectations not only solves the non-fundamentalness problem created by foresight but also makes it possible to identify structural shocks through an appropriate expectations-based scheme. In particular, when expectations-based identifying assumptions are com-
bined with standard short-run exogeneity restrictions, the expectations-based approach is found to be useful for the identification of surprise spending shocks.

The application of the approach to U.S. data supports concerns raised by Leeper, Walker, and Yang (2011) and Ramey (2011b) on the validity of the results of previous empirical studies due to the influence of policy foresight. The expectations-based approach indicates positive short-run output effects of federal government expenditures, but negative medium-term effects due to falling consumption and investment. The standard SVAR identification scheme, on the other hand, predicts stronger and more persistent effects on consumption, investment, and output. However, Granger causality tests suggest that, unlike the surprise shocks identified by the expectations-based scheme, the shocks identified by the standard scheme are forecastable.

In addition to policy foresight, several alternative explanations for the differences to previous studies are conceivable. In particular, the post-1980 period is often argued to have smaller average fiscal multipliers than the pre-1980 period. In addition, the structure of government spending is likely to matter, given that more than 70 percent of U.S. federal spending falls on defense-related expenditures. An investigation of the effects of other types of expenditures such as state and local government spending, for which expectations data is also available from the Survey of Professional Forecasters, would thus be a useful extension of the analysis conducted in this chapter.

### 3.A Analytical solution

This appendix provides a detailed derivation of the analytical solution of the model. The steps are as follows. First, the non-stochastic steady state solution is derived from the non-linear equilibrium conditions. The equilibrium system is then log-linearized around the non-stochastic steady state and reduced to a two-dimensional first-order linear difference equation in capital and consumption, given the stochastic processes for TFP and government spending. Finally, the parameters in the recursive laws of motion for capital and consumption are derived by the method of undetermined coefficients (see Uhlig, 1999).

---

\(^{18}\) See, for instance, Bilbiie et al. (2008) and Blanchard and Perotti (2002).
Non-linear equilibrium. The non-linear equilibrium conditions are as follows:

Labor/leisure: \( Ac_t = (1 - \alpha) y_t / n_t \),

Euler equation: \( E_t = \beta E_{t+1} / c_{t+1} \),

Real return: \( R_t = 1 - \delta + \alpha y_t / k_{t-1} \),

Production: \( y_t = a_t k_{t-1}^\alpha n_t^{1-\alpha} \),

Feasibility: \( y_t = c_t + k_t - (1 - \delta) k_{t-1} + g_t \),

TFP: \( \log a_t = \rho_a \log a_{t-1} + \varepsilon_{a,t} \),

Gov. expenditures: \( \log \left( \frac{g_t}{\bar{g}} \right) = \rho_g \log \left( \frac{g_{t-1}}{\bar{g}} \right) + \varepsilon_{g,t}^a + \varepsilon_{a,g,t} \),

where investment has been eliminated in the feasibility constraint using the capital accumulation equation.

Steady state. Let variables without time subscript denote non-stochastic steady state values. The TFP process implies \( a = 1 \) since \( \rho_a \in (0, 1] \). The Euler equation yields \( R = \beta^{-1} \). The real return equation can be solved for the output-to-capital ratio:

\[
\frac{y}{k} = \frac{R - 1 + \delta}{\alpha} = \frac{\beta^{-1} - 1 + \delta}{\alpha}.
\]

The production function implies that

\[
\frac{y}{n} = \frac{k^{\alpha} n^{1-\alpha}}{n} = \left( \frac{k}{n} \right)^\alpha, \quad \frac{y}{k} = \frac{k^{\alpha} n^{1-\alpha}}{k} = \left( \frac{k}{n} \right)^{\alpha-1}.
\]

From the second equation, we have \( k/n = (y/k)^{(1/(\alpha-1))} = (y/k)^{-1/(1-\alpha)} \). Substituting this expression into the first equation yields an expression for the output-to-labor ratio:

\[
\frac{y}{n} = \left( \frac{y}{k} \right)^{-\frac{\alpha}{1-\alpha}} = \left( \frac{\beta^{-1} - 1 + \delta}{\alpha} \right)^{-\frac{\alpha}{1-\alpha}}.
\]

The labor/leisure tradeoff then yields \( c = A^{-1}(1 - \alpha) y/n \). Dividing the feasibility constraint by \( y \) and re-writing yields

\[
n = \frac{c(y/n)^{-1}}{1 - \delta k/y - g/y}.
\]
Taking \( s_g = g/y \) as given, the government spending equation implies \( g = \bar{g} \), since \( \rho_g \in [0, 1) \). The remaining steady state solutions follow as

\[
y = (y/n)n, \quad k = (k/y)y, \quad g = \bar{g} = s_g y.
\]

**Log-linearized system.** The log-linearized system is given by

- **Labor/leisure:** \( \hat{n}_t = \hat{n}_t - \hat{c}_t \), \( t = 3.19 \)
- **Euler equation:** \( 0 = E_t[\hat{c}_t - \hat{c}_{t+1} + \hat{R}_{t+1}] \), \( t = 3.20 \)
- **Production:** \( \hat{y}_t = \hat{a}_t + \alpha \hat{k}_{t-1} + (1 - \alpha) \hat{n}_t \), \( t = 3.21 \)
- **Feasibility:** \( \alpha \hat{c}_t = y\hat{y}_t - k\hat{k}_t + (1 - \delta)k\hat{k}_{t-1} - gg \), \( t = 3.22 \)
- **Real return:** \( \hat{R}_t = \frac{\alpha y}{Rk}(\hat{y}_t - \hat{k}_{t-1}) \), \( t = 3.23 \)
- **TFP:** \( \hat{a}_t = \rho_a \hat{a}_{t-1} + \varepsilon_{a,t} \), \( t = 3.24 \)
- **Gov. expenditures:** \( \hat{g}_t = \rho_g \hat{g}_{t-1} + \varepsilon_{g,t} + \varepsilon_{g,t-2} \), \( t = 3.25 \)

where \( \hat{x}_t = \log(x_t/x) \) such that \( x_t = x \exp(\hat{x}_t) \approx x(1 + \hat{x}_t) \) for \( \hat{x}_t \approx 0 \).

**Difference equations.** The log-linearized system is now reduced to the two first-order linear difference equations reported in the main text. Substituting (3.19) into (3.21), leading the result by one period, and re-arranging terms yields

\[
\hat{y}_{t+1} = \frac{\hat{a}_{t+1} + \alpha \hat{k}_t - (1 - \alpha)\hat{c}_{t+1}}{\alpha}.
\]

Substituting (3.23) into (3.20) yields

\[
0 = E_t \left[ \hat{c}_t - \hat{c}_{t+1} + \frac{\alpha y}{Rk}(\hat{y}_{t+1} - \hat{k}_t) \right].
\]

Combining the latter two expressions gives the first difference equation:

\[
0 = E_t \left[ \hat{c}_t - \hat{c}_{t+1} + \frac{\alpha y}{Rk} \left( \frac{1}{\alpha} \hat{a}_{t+1} + \hat{k}_t - \frac{1 - \alpha}{\alpha} \hat{c}_{t+1} - \hat{k}_t \right) \right],
\]

\[
= E_t \left[ \hat{c}_t - \left( 1 + \frac{y}{k} \frac{1 - \alpha}{R} \right) \hat{c}_{t+1} + \frac{y}{k R} \hat{a}_{t+1} \right].
\]
Using (3.21) in (3.22) yields
\[ cc_t = y \hat{a}_t + [\alpha y + (1 - \delta)k] \hat{k}_{t-1} + (1 - \alpha) y \hat{a}_t - k \hat{k}_t - g \hat{g}_t. \]

Substituting out (3.19) in the last expression, using (3.26) lagged by one period, re-
arranging terms, and taking expectations yields the second difference equation:
\[ 0 = E_t \left[ \left( c + \frac{1 - \alpha}{\alpha} \right) \hat{c}_t + k \hat{k}_t - \frac{y}{\alpha} \hat{a}_t - (y + (1 - \delta)k) \hat{k}_{t-1} + g \hat{g}_t \right]. \]

The reduced system is thus given by
\[ 0 = E_t [\hat{c}_t - \phi_1 \hat{c}_{t+1} + \phi_2 \hat{a}_{t+1}], \quad (3.27) \]
\[ 0 = E_t [\phi_3 \hat{c}_t + \phi_4 \hat{k}_t - \phi_5 \hat{a}_t - \phi_6 \hat{k}_{t-1} + \phi_7 \hat{g}_t], \quad (3.28) \]

where \( \phi_1 = 1 + R^{-1}(1 - \alpha)y/k, \phi_2 = R^{-1}y/k, \phi_3 = c + y(1 - \alpha)/\alpha, \phi_4 = k, \phi_5 = y/\alpha, \phi_6 = y + (1 - \delta)k \) and \( \phi_7 = g \).

**Recursive laws of motion.** Next, guess the recursive laws of motion
\[ \hat{k}_t = \eta_{kk} \hat{k}_{t-1} + \eta_{ka} \hat{a}_t + \eta_{kg} \hat{g}_t + \eta_{k\varepsilon,1} \epsilon^a_{g,t} + \eta_{k\varepsilon,2} \epsilon^a_{g,t-1}, \]
\[ \hat{c}_t = \eta_{ck} \hat{k}_{t-1} + \eta_{ca} \hat{a}_t + \eta_{cg} \hat{g}_t + \eta_{c\varepsilon,1} \epsilon^a_{g,t} + \eta_{c\varepsilon,2} \epsilon^a_{g,t-1}. \]

Repeatedly substituting the latter into (3.27) and (3.28) and using \( E_t \hat{g}_{t+1} = \rho_g \hat{g}_t + \epsilon^a_{g,t+1} \)
and \( E_t \hat{a}_{t+1} = \rho_a \hat{a}_t \) yields, after some tedious but straightforward algebra:
\[ 0 = \left[ (1 - \phi_1 \eta_{kk}) \eta_{ck} \right] \hat{k}_{t-1} + \left[ \eta_{ca} (1 - \phi_1 \rho_a) + \phi_2 \rho_a - \phi_1 \eta_{ck} \eta_{ka} \right] \hat{a}_t + \left[ \eta_{cg} (1 - \phi_1 \rho_g) - \phi_1 \eta_{ck} \eta_{kg} \right] \hat{g}_t + \left[ \eta_{c\varepsilon,1} - \phi_1 (\eta_{c\varepsilon,2} + \eta_{ck} \eta_{ke,1}) \right] \epsilon^a_{g,t} + \left[ \eta_{c\varepsilon,2} - \phi_1 (\eta_{cg} + \eta_{ck} \eta_{ke,2}) \right] \epsilon^a_{g,t-1}, \quad (3.29) \]
\[ 0 = \left[ \phi_3 \eta_{ck} + \phi_4 \eta_{kk} - \phi_6 \right] \hat{k}_{t-1} + \left[ \phi_3 \eta_{ca} + \phi_4 \eta_{ka} - \phi_5 \right] \hat{a}_t + \left[ \phi_3 \eta_{cg} + \phi_4 \eta_{kg} + \phi_7 \right] \hat{g}_t + \left[ \phi_3 \eta_{c\varepsilon,1} + \phi_4 \eta_{ke,1} \right] \epsilon^a_{g,t} + \left[ \phi_3 \eta_{c\varepsilon,2} + \phi_4 \eta_{ke,2} \right] \epsilon^a_{g,t-1}. \]
Solving for the dynamics. Finally, one can solve for the coefficients \( \eta \) in the recursive laws of motion. Both of the above equations must hold with equality for all values of the state variables. First, set \( \hat{a}_t = \hat{g}_t = \epsilon_{a,t}^a = \epsilon_{g,t}^a = 0 \):

\[
0 = [(1 - \phi_1 \eta_{kk}) \eta_{ck}] \hat{k}_{t-1}, \\
0 = [\phi_3 \eta_{ck} + \phi_4 \eta_{kk} - \phi_6] \hat{k}_{t-1}.
\]

Since both equations also need to hold for any value of \( \hat{k}_{t-1} \), it must be that

\[
0 = (1 - \phi_1 \eta_{kk}) \eta_{ck}, \\
0 = \phi_3 \eta_{ck} + \phi_4 \eta_{kk} - \phi_6.
\]

The second equation implies

\[
\eta_{ck} = \frac{\phi_6}{\phi_3} - \frac{\phi_4}{\phi_3} \eta_{kk},
\]

and the first equation implies

\[
0 = \phi_1 \phi_3 \eta_{kk}^2 - (\phi_1 \phi_6 + \phi_4) \eta_{kk} + \phi_6.
\]

with solutions

\[
\eta_{kk}^\pm = \frac{1}{2} \pm \sqrt{\left(\frac{\phi_1^{-1} + \phi_6/\phi_4}{2}\right)^2 - \frac{\phi_6}{\phi_3 \phi_4}}.
\]

Similarly, comparing coefficients on \( \hat{a}_t \) gives

\[
\eta_{ka} = \frac{\phi_5}{\phi_4} - \frac{\phi_3}{\phi_4} \eta_{ca}, \\
\eta_{ca} = \frac{\phi_1 (\phi_5/\phi_4) \eta_{ck} - \phi_2 \rho_a}{1 - \phi_1 \rho_a + \phi_1 (\phi_3/\phi_4) \eta_{ck}}.
\]

Comparing coefficients on \( \hat{g}_t \) yields

\[
\eta_{kg} = -\left(\frac{\phi_7}{\phi_4} + \frac{\phi_3}{\phi_4} \rho_g\right), \\
\eta_{cg} = \frac{-\phi_1 (\phi_7/\phi_4) \eta_{ck}}{1 + \phi_1 [(\phi_3/\phi_4) \eta_{ck} - \rho_g]}.
\]
Further, comparing coefficients on $\varepsilon_{g,t-1}$ gives

$$\eta_{cz,2} = -\frac{\phi_4}{\phi_3} \eta_{kz,2}, \quad \eta_{kz,2} = -\frac{\eta_{cg}}{\eta_{ck} + \phi_4/(\phi_1\phi_3)}.$$  

Finally, comparing coefficients on $\varepsilon_{g,t}$ yields

$$\eta_{kz,1} = -\frac{\phi_4}{\phi_3} \eta_{kz,1}, \quad \eta_{kz,1} = -\frac{\eta_{cz,2}}{\eta_{ck} + \phi_4/(\phi_1\phi_3)}.$$  

**Modifications.** When spending reacts to lagged TFP, the coefficient on $\hat{a}_t$ in (3.29) changes to $\eta_{ca}(1 - \phi_1\rho_a) + \phi_2\rho_a - \phi_1(\eta_{ck}\eta_{ka} + \eta_{cg}\rho_{ga})$. Thus, the only coefficients in the recursive laws of motion that are affected are $\eta_{ca}$ and $\eta_{ka}$. For $\rho_{ga} \in \mathbb{R}$, they are

$$\eta^*_ka = \frac{\phi_5}{\phi_4} - \frac{\phi_3}{\phi_4} \eta^*_ca, \quad \eta^*_ca = \frac{\phi_1(\phi_5/\phi_4)\eta_{ck} - \phi_2\rho_a + \rho_{ga}\phi_1\eta_{cg}}{1 - \phi_1\rho_a + \phi_1(\phi_3/\phi_4)\eta_{ck}}.$$  

If $\rho_{ga} = 0$, it follows that $\eta^*_ka = \eta_{ka}$ and $\eta^*_ca = \eta_{ca}$.  

When spending reacts to current TFP, the coefficient on $\hat{a}_t$ in (3.29) changes to $\eta_{ca}(1 - \phi_1\rho_a) + \phi_2\rho_a - \phi_1(\eta_{ck}\eta_{ka} + \eta_{cg}\rho_{ga}\rho_a)$. In this case,

$$\eta^{**}ka = \frac{\phi_5}{\phi_4} - \frac{\phi_3}{\phi_4} \eta^{**}ca, \quad \eta^{**}ca = \frac{\phi_1(\phi_5/\phi_4)\eta_{ck} - \phi_2\rho_a + \rho_{ga}\rho_a\phi_1\eta_{cg}}{1 - \phi_1\rho_a + \phi_1(\phi_3/\phi_4)\eta_{ck}}.$$  

If $\rho_{ga} = 0$, it follows that $\eta^{**}ka = \eta^*_ka = \eta_{ka}$ and $\eta^{**}ca = \eta^*_ca = \eta_{ca}$.  

### 3.B Detailed data description

This appendix provides details on data sources and data definitions. Throughout, NIPA refers to the National Income and Product Accounts of the Bureau of Economic Analysis, BLS to the Bureau of Labor Statistics, ALFRED to the Archival Federal Reserve Economic Data of the Federal Reserve Bank of St. Louis, and SPF to the Survey of Professional Forecasters of the Federal Reserve Bank of Philadelphia. All time series are provided in seasonally adjusted terms from the original sources, except the data on federal debt and the T-bill rate which are not seasonally adjusted.  

- **Government spending, realization:** Real federal government consumption plus...
gross investment; the nominal series is taken from NIPA Table 1.1.5. Line 22; the series is then scaled to constant 2005 prices by its deflator (NIPA Table 1.1.4. Line 22) and converted into natural logarithms.

• *Government spending, forecasts:* One to three-quarter ahead forecasts of real federal government consumption plus gross investment; the level series is the mean prediction of SPF variable RFEDGOV; given breaks in levels due to NIPA base year changes, the forecasts are scaled to constant 2005 prices by backcasting the actual growth rates and imposing the average growth rate over the sample at the break points.

• *Government spending, expectational error:* First difference of natural logarithms of realized spending and the prediction thereof made one quarter earlier.

• *Output:* Real gross domestic product; the nominal series is taken from NIPA Table 1.1.5. Line 1; the series is then scaled to constant 2005 prices by its deflator (NIPA Table 1.1.4. Line 1) and converted into natural logarithms.

• *Consumption:* Real personal consumption expenditure; the nominal series is taken from NIPA Table 1.1.5. Line 2; the series is then scaled to constant 2005 prices by its deflator (NIPA Table 1.1.4. Line 2) and converted into natural logarithms.

• *Investment:* Real gross private investment; the nominal series is taken from NIPA Table 1.1.5. Line 7; the series is then scaled to constant 2005 prices by its deflator (NIPA Table 1.1.4. Line 7) and converted into natural logarithms.

• *Real wage:* Real hourly compensation in the business sector; BLS series ID: PRS84006153; the original series is converted into natural logarithms.

• *Interest rate:* 3-month treasury bill rate, secondary market rate; series TB3MS in ALFRED database; the interest rate is expressed in annual terms.

• *Debt-to-GDP ratio:* total end-of-period federal government debt divided by nominal GDP; public debt data: series GFDEBTN in ALFRED database.
Chapter 4

Financial Stress, Government Policies, and the Consequences of Deficit Financing*

Abstract

We study the effects of deficit-financed government policies during a financial crisis based on a structural macroeconomic model with financial intermediation. Related studies that consider financial intermediation in a structural framework take for granted that the government does not rely upon intermediary financing. We set up a model where intermediaries not only channel funds from households to firms but also to the government. Accordingly, subject to balance sheet constraints, intermediaries use deposits to purchase government bonds and to provide loans to firms. Our framework thereby extends existing frameworks by allowing for different types of assets in intermediary portfolios. The latter has important consequences for the effectiveness of government policies. In particular, our results highlight a crowding out mechanism that has government borrowing tighten intermediary constraints such that a fiscal expansion can further raise credit spreads during a crisis, which can even further reduce economic activity.

4.1 Introduction

The fiscal response to the recent financial and economic crisis took the form of financial sector support measures and economic stimulus packages that were financed through budgetary deficits. What are the effects of such policies in a situation of financial

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stress such as the recent crisis? Standard macroeconomic models are not set up for policy analysis during crises, but new types of models have been developed that can be used to answer this question (see Gertler and Karadi, 2011; Gertler and Kiyotaki, 2010). However, these models take for granted that government policies are fully funded through markets that are not part of the relevant financial system. As we discuss below, this assumption seems unrealistic in view of the actual practice.

This chapter develops a structural macroeconomic model that integrates government deficit financing and financial intermediation with frictions in the intermediation process. These frictions imply a financial accelerator that is able to generate a deep financial crisis following a deterioration of intermediary balance sheets. We use this model to study the effects of deficit-financed policies during this type of crisis.

Our framework includes financial intermediaries that channel funds or deposits from households (the saving agents) to non-financial firms and the government (the borrowing agents). The intermediation process is subject to a similar agency problem between depositors and intermediaries as in Gertler and Karadi (2011) and Gertler and Kiyotaki (2010). As those studies show, the latter leads to endogenous balance sheet constraints as key elements of a powerful financial accelerator mechanism. This accelerator mechanism generates dynamics which broadly reflect the relevant economic dynamics of a financial crisis. However, unlike the models in those studies, our framework allows for different classes of assets in intermediary portfolios instead of only one class.

The specific set-up is as follows. The intermediary asset portfolio consists of government bonds next to private claims from loans to non-financial firms. The portfolio size is tied to intermediary equity capital through endogenous leverage constraints, but the intermediaries can shift the composition of their portfolios towards assets with higher expected returns. Through this mechanism, the expected returns on bonds and private claims are jointly determined in equilibrium. Accordingly, when a new equilibrium arises in either the bond market or the market for private claims, everything else equal the expected return on bonds tends to deviate from the expected return on private claims. Individual intermediaries seek to exploit the differences in expected returns by altering the composition of their portfolios. In the general equilibrium, such arbitrage behavior by financial intermediaries leads to co-movements between different
credit spreads relative to the rates at which intermediaries obtain funding.

We use this model to highlight the link between government policies and deficit financing in a situation of financial stress. In particular, we analyze the effects of demand (i.e. spending) stimulus and measures targeting the financial sector (i.e. transfers to intermediaries, zero interest loans, and loans at penalty rates). This set of policies is sufficient to explain the key implications of the model, but it also suitably captures the main fiscal policy measures that were applied during the recent crisis. The policies are financed by issuing bonds to intermediaries or by raising lump-sum taxes directly from households. The latter allows the government to circumvent the financial frictions and serves as a benchmark for our results.

Our findings suggest that intermediary financing has important consequences for the effectiveness of government policies. An early demand stimulus dampens the recession due to a financial crisis for some time (i.e. it reduces output losses). However, the stimulus tends to prolong the downturn later on. Moreover, the announcement of a stimulus has the potential to deepen the crisis until the stimulus is actually implemented. We also find that financial sector policies become less effective under intermediary financing. Some policies such as loans with relatively early repayment can also deepen the downturn. Temporary support can however bring initial stabilization gains if the cost to intermediaries is shifted towards later periods. Overall, these findings have implications beyond the predictions of previous studies that are apt to overturn conclusions on the effectiveness of deficit-financed government policies.

Key to understand our findings are the effects of government borrowing on intermediary balance sheet constraints and the associated adverse impacts on the cost of credit to non-financial firms. In our model, a fiscal expansion is associated with an economy-wide increase in credit spreads, as higher government deficits tighten intermediary balance sheet constraints. The rise in spreads lowers non-financial sector investment, which can offset the output gain of a demand stimulus. The same mechanism also reduces the effectiveness of financial sector policies. The fact that intermediary balance sheet constraints are forward-looking explains the links between the timing and the effects of government policies.

This chapter is closely related to the above-mentioned studies that emerged out
of the experience of the recent crisis.\footnote{Other related studies include, for instance, Angeloni and Faia (2010), Bean, Paustian, Penalver, and Taylor (2010), Christiano and Ikeda (2011), and Gertler, Kiyotaki, and Queralto (2010).} Gertler and Karadi (2011) evaluate the effects of government (central bank) credit intermediation, financed by issuing government debt to households, to offset a disruption of private financial intermediation. Gertler and Kiyotaki (2010) consider a generalization of the model in Gertler and Karadi (2011) with an interbank market and also analyze the effects of government equity injections, financed by raising lump-sum taxes from households. Thus, in these studies, government policies are financed directly by households. Our model allows for deficit financing of fiscal policy through financial intermediaries to do justice to the actual practice of fiscal financing in developed countries.

In fact, many financial institutions in developed countries are active in government funding markets. In the euro area, a significant fraction of monetary financial institutions’ assets consists of government securities and direct loans to the government: on average about 9\% of total assets and about 58\% of the value of loans to non-financial corporations.\footnote{Source: European Central Bank; aggregated balance sheet of euro area monetary financial institutions, March 2011; \url{http://www.ecb.int/stats/money/aggregates/bsheets}.} In addition, EU banks hold primarily domestic government securities (see ECB, 2010). Using a closed-economy approach, as we do, thus seems sufficient to capture the key elements of sovereign funding structures in Europe. Outside of Europe, Japanese bank holdings of government securities as a proportion of total assets have recently gone up to an all-time high, as banks have become the dominant buyers of government bonds. In the UK and the U.S., domestic depository institutions’ claims on the government amount to approximately 6\% and 8\% of GDP, respectively (see IMF, 2010c). Hence, government securities holdings by domestic financial institutions play an important role in most high-income countries.

Overall, we therefore view this chapter as a further step towards a more realistic description of financial markets and as one of the first steps to reflect fiscal-financial linkages in macroeconomic models. Recurring post-crisis concerns on the sustainability of government debt in developed countries and the associated spillover effects across financial systems suggest that these are steps into a relevant direction.

The remainder of the chapter is structured as follows. Section 4.2 lays out the model.
Section 4.3 discusses the results of model-based simulations. It first compares the effects of deficit-financed and lump-sum-tax-financed changes in government purchases in comparison to a baseline model without financial intermediation to explain the main mechanisms at play. It then analyzes the effects of alternative fiscal policy responses to a simulated crisis to investigate the stabilization properties of different policies. Section 4.4 briefly reviews the related empirical literature to connect the key mechanisms and predictions of the model to the available evidence. Section 4.5 concludes.

4.2 Model description

We describe a monetary model with sticky prices and financial intermediation that builds on Christiano, Eichenbaum, and Evans (2005) and Gertler and Karadi (2011).

The model has a private sector and a public sector. The private sector consists of a non-financial sector that is formed by households and firms, and a financial sector that is formed by financial intermediaries.

The firm production chain is as follows. Capital producers combine used capital purchased from intermediate goods producers with investment goods to produce new productive capital which is again purchased by intermediate goods producers. The latter rent labor services from households and issue claims to financial intermediaries to finance their capital acquisition. They produce differentiated goods that are bought, re-packaged, and sold by retail firms in a monopolistically competitive market. Final goods producers buy those goods and combine them into a single output good.

The public sector is formed by a monetary authority that sets the risk-free nominal interest rate and a government that conducts purchases of the final good and financial sector policies. The government finances its operations by issuing debt to financial intermediaries or by raising lump-sum taxes from households. The intermediaries take funds from depositors which are remunerated at the risk-free nominal interest rate.

4.2.1 Households

There is a continuum of infinitely lived households with identical preferences and identical asset endowments. Following Gertler and Karadi (2011), within each household
there is a fraction \( 1 - \zeta \) of workers that supply labor to firms and a fraction \( \zeta \) of bankers that operate financial intermediaries. There is perfect consumption insurance within the family. Households save by holding deposits at intermediaries which they do not own. Financial intermediaries have finite life times, to exclude the self-financing equilibrium. Thus, at the beginning of each period, with probability \( 1 - \theta \) an individual intermediary exits and with probability \( \theta \) the intermediary continues operating. If the intermediary exits, the respective bankers become workers and transfer any retained capital back to the household which owns that intermediary. Thus every period \((1 - \theta)\zeta\) bankers become workers. To keep the relative proportions fixed, a similar number of workers become bankers. New bankers receive a start-up transfer from their household, as described below.

Household preferences depend on consumption and labor supply, with habit formation in consumption as in Christiano et al. (2005) in order to capture consumption dynamics. The objective of a representative household in period \( t \) is to maximize expected discounted utility

\[
E_t \sum_{s=0}^{\infty} \beta^s \left[ \log(c_{t+s} - \nu c_{t-1+s}) - (1 + \varphi)^{-1} h_{t+s}^{1+\varphi} \right], \quad \beta \in (0, 1), \quad \nu \in [0, 1), \quad \varphi \geq 0,
\]

subject to the period-by-period budget constraint

\[
c_t + d_t + \tau_t \leq w_t h_t + (1 + r_t^d) d_{t-1} + \Sigma_t,
\]

where \( c_t \) denotes consumption of final goods, \( h_t \) denotes hours worked, \( w_t \) is the hourly wage rate, \( d_{t-1} \) are beginning-of-period deposits, \( d_t \) are end-of-period deposits, \( r_t^d \) is the net real interest rates on deposits, \( \tau_t \) are lump-sum tax payments, and \( \Sigma_t \) collects payouts from ownership of both non-financial and financial firms, net of transfers given to household members that enter as bankers at time \( t \).

The household’s decision problem is subject to a no-Ponzi game condition, and the household takes \( w_t, r_t^d, \tau_t, \Sigma_t, \) prices, and its initial wealth endowment \( d_{-1} \) as given. The first-order conditions corresponding to the solution of the household’s problem are

\[3\]Throughout, real (nominal) variables are denoted by lower (capital) letters, and variables without time subscript denote non-stochastic steady state values.
given by

\begin{align*}
    c_t & : \quad \lambda_t = (c_t - \nu c_{t-1})^{-1} - \beta \nu (E_t c_{t+1} - \nu c_t)^{-1}, \\
    h_t & : \quad h_t^i = \lambda_t w_t, \\
    d_t & : \quad 1 = \beta E_t \Lambda_{t,t+1}(1 + r^d_{t+1}),
\end{align*}

(4.1) (4.2) (4.3)

where \( \lambda_t \) denotes the Lagrangian multiplier associated with the budget constraint and \( \Lambda_{t,t+s} = \lambda_{t+s}/\lambda_t \) for \( s \geq 0 \). As \( \lambda_t > 0 \), the budget constraint holds with equality.

### 4.2.2 Financial intermediaries

Financial intermediaries are competitive and located on a continuum indexed by \( j \in [0, 1] \). The intermediaries use the deposits obtained from households to purchase claims issued by intermediate goods firms and government bonds. Intermediaries thus act as specialists that assist in channeling funds from agents with a surplus of funds to agents with deficits of funds, where the latter include the government. The need for the government to resort to intermediaries is motivated by size arguments: bond issuance typically occurs in large tranches that cannot be handled by small investors.

To introduce different assets, we characterize the intermediary problem as a two-stage procedure that seems to have sufficient practical appeal. In particular, we assume that each intermediary is operated by a bank manager (or bank board) who makes size decisions and a portfolio manager (or portfolio department) who decides on the structure of assets. In the first stage, the bank manager chooses the total amount of assets relative to deposits to maximize the expected transfer to the household that owns the respective intermediary, similarly as in Gertler and Karadi (2011). Also following the latter, a moral hazard problem constrains the bank manager’s ability to obtain funds. In the second stage of the intermediary problem, for a given portfolio size, the portfolio manager chooses portfolio weights to maximize the same objective as the bank manager.

Total assets of intermediary \( j \) at the end of period \( t \) are given by

\[ P_{j,t} = q_t s_{j,t}^k + s_{j,t}^h, \]
where \( s_{j,t}^k \) denote claims on intermediate goods firms by intermediary \( j \) that have the relative price \( q_t \) and that pay a net real return \( r_{t+1}^k \) at the beginning of period \( t+1 \), and \( s_{j,t}^b \) are intermediary \( j \)’s government bond holdings that pay a net real return \( r_{t+1}^b \) at the beginning of period \( t+1 \). The balance sheet of intermediary \( j \) thus looks as follows:

\[ p_{j,t} = d_{j,t} + n_{j,t}, \]

where \( d_{j,t} \) denote deposits by households at intermediary \( j \) and \( n_{j,t} \) denotes the intermediary’s net worth. The latter evolves over time as the difference between earnings on assets and interest payments on liabilities minus payments or costs due to portfolio adjustments:

\[ n_{j,t+1} = (1 + r_{t+1}^p) p_{j,t} - (1 + r_{t+1}^d) d_{j,t} - \Omega(\omega_{j,t}) n_{j,t}, \]

where \( r_{t+1}^p \) is the net real portfolio return. We further define portfolio weights \( \omega_{j,t} = q_t s_{j,t}^k / p_{j,t} \) and \( 1 - \omega_{j,t} = s_{j,t}^b / p_{j,t} \), such that the ex-post gross portfolio return satisfies

\[ 1 + r_{t}^p = (1 + r_{t-1}^k) \omega_{j,t-1} + (1 + r_{t}^b)(1 - \omega_{j,t-1}). \tag{4.4} \]

The term \( \Omega(\omega_{j,t}) n_{j,t} \) above measures convex portfolio adjustment costs that are scaled by the level of net worth. We introduce these costs to achieve stationarity, to be able to use standard local approximation techniques. Such costs could come, for instance, from fees that are incurred when assets are bought and sold on the market.\(^4\)

In the context of our model, those fees are eventually paid out to households. The costs are scaled by the level of net worth to allow for aggregation, as conducted below, motivated by the idea that the total costs that an individual intermediary incurs on portfolio changes should depend on the total scale of that intermediary’s operations. We apply the following functional form:

\[ \Omega(x) = \frac{\varpi}{2} (x - \bar{\omega})^2, \quad \Omega'(x) = \varpi(x - \bar{\omega}), \quad \varpi > 0, \quad \bar{\omega} \in (0,1). \]

The adjustment costs are thus increasing in deviations of the portfolio weight \( \omega_{j,t} \) from

\(^4\)The existence of costly portfolio adjustments is supported by aggregate estimates and micro evidence of infrequent portfolio changes by U.S. stockholders (see Luttmer, 1999; Bonaparte and Cooper, 2010).
a long-run target $\bar{\omega}$. The latter pins down the steady state portfolio weights and thus helps to match steady state supply of government bonds and the steady state level of private assets in the general equilibrium, as shown in Appendix 4.A.

**Bank manager**

At the beginning of period $t + 1$, after financial payouts have been made, an individual financial intermediary continues operating with probability $\theta$ and exits with probability $1 - \theta$, in which case it transfers its retained capital to its household. The bank manager’s objective in period $t$ is therefore to maximize expected terminal wealth, given by

$$V_{j,t} = E_t \sum_{i=0}^{\infty} (1 - \theta)\theta^i \beta^{i+1} \Lambda_{t,t+1+i} n_{j,t+1+i}. $$

However, following Gertler and Karadi (2011), a costly enforcement problem constrains the ability of financial intermediaries to obtain funds from depositors. In particular, at the beginning of period $t$, before financial payouts are made, the bank manager can divert a fraction $\lambda$ of total assets. The depositors can then force the intermediary into bankruptcy and recover the remaining assets, but it is too costly for the depositors to recover the funds that the banker diverted. Accordingly, for the depositors to be willing to supply funds, the incentive constraint $V_{j,t} \geq \lambda p_{j,t}$ must be satisfied. That is, the opportunity cost to the banker of diverting assets cannot be smaller than the gain from diverting assets. It can be shown that $V_{j,t}$ can be expressed as follows:

$$V_{j,t} = v_t p_{j,t} - \eta_t d_{j,t} - \varrho_t n_{j,t},$$

$$V_t = \beta E_t \Lambda_{t,t+1} \{ (1 - \theta)(1 + r_{t+1}^p) + \theta x_{t,t+1} v_{t+1} \}, \quad x_{t,t+1} = p_{j,t+1}/p_{j,t}, \quad (4.5)$$

$$\eta_t = \beta E_t \Lambda_{t,t+1} \{ (1 - \theta)(1 + r_{t+1}^d) + \theta z_{t,t+1} \eta_{t+1} \}, \quad z_{t,t+1} = d_{j,t+1}/d_{j,t}, \quad (4.6)$$

$$\varrho_t = \beta E_t \Lambda_{t,t+1} \{ (1 - \theta)\Omega(\omega_{j,t}) + \theta f_{t,t+1} \varrho_{t+1} \}, \quad f_{t,t+1} = n_{j,t+1}/n_{j,t}. \quad (4.7)$$

Holding the other variables constant, the variable $v_t$ is the expected discounted marginal gain of an additional unit of assets. The variable $\eta_t$ is expected discounted marginal cost of another unit of deposits. The variable $\varrho_t$ is the expected discounted marginal cost of another unit of net worth conditional on portfolio changes.
We assume that the bank manager takes the expected returns and the portfolio weights as given when deciding on the total size of assets. The Lagrangian of the bank manager’s optimization problem is given by \( L = V_{j,t} + \mu_t(V_{j,t} - \lambda p_{j,t}) \), where \( \mu_t \geq 0 \) is the Lagrangian multiplier associated with the incentive constraint. The first-order conditions are

\[
p_{j,t} : (1 + \mu_t)(v_t - \eta_t) - \mu_t \lambda = 0,
\]

\[
\mu_t : (v_t - \eta_t - \lambda)p_{j,t} + (\eta_t - \varrho_t)n_{j,t} \geq 0.
\]

The last condition holds with equality if \( \mu_t > 0 \), otherwise it holds with strict inequality. The condition for \( p_{j,t} \) yields \( \mu_t = \left[ \frac{\lambda}{(v_t - \eta_t) - 1} \right]^{-1} \). The multiplier is therefore strictly positive if \( \lambda > v_t - \eta_t \). That is, the incentive constraint holds with equality if the banker has an incentive to divert funds obtained from depositors and go bankrupt instead of continuing to operate with the additional funds. We assume that the incentive constraint binds within a local region of the non-stochastic steady state and verify that it does bind in the steady state in the calculations in Appendix 4.A.

In the optimum, the total amount of intermediary assets is then tied to intermediary net worth through the leverage constraint \( p_{j,t} = \phi_t n_{j,t} \), where

\[
\phi_t = \frac{\eta_t - \varrho_t}{\lambda - (v_t - \eta_t)} \tag{4.8}
\]

denotes the intermediary’s leverage ratio of assets over net worth. As indicated by (4.8), a higher marginal gain from increasing assets \( v_t \) supports a higher leverage ratio in the optimum. A higher marginal cost of deposits \( \eta_t \) lowers the leverage ratio. Higher marginal adjustment costs \( \varrho_t \) and a larger fraction of divertable funds \( \lambda \) also lower the leverage ratio.

**Portfolio manager**

The portfolio manager determines the asset structure of intermediary \( j \)'s balance sheet by choosing portfolio weights to maximize the same objective as the bank manager, taking as given the total size of assets \( p_{j,t} \) and the interest rates \( r_{i,t+1}^i, i = k, b, d. \)
Using the portfolio weights, the holdings of individual assets by intermediary \( j \) satisfy \( q_t^k s_{j,t}^k = \omega_{j,t} p_{j,t} \) and \( s_{j,t}^b = (1 - \omega_{j,t}) p_{j,t} \). Net worth of intermediary \( j \) can therefore be re-written as follows:

\[
\begin{align*}
n_{j,t+1} & = (1 + r_{t+1}^k) q_t^k s_{j,t}^k + (1 + r_{t+1}^b) s_{j,t}^b - (1 + r_{t+1}^d) d_{j,t} - \Omega(\omega_{j,t}) n_{j,t}, \\
& = (r_{t+1}^k - r_{t+1}^d) \omega_{j,t} p_{j,t} + (r_{t+1}^b - r_{t+1}^d)(1 - \omega_{j,t}) p_{j,t} + [1 + r_{t+1}^d - \Omega(\omega_{j,t})] n_{j,t}.
\end{align*}
\]

The optimization problem of the portfolio manager is then given by

\[
\max_{\omega_{j,t}} E_t \sum_{i=0}^{\infty} (1 - \theta)^{\theta^i} \beta^{i+1} A_{t,t+1+i} n_{j,t+1+i}.
\]

The first-order condition looks as follows:

\[
E_t (r_{t+1}^k - r_{t+1}^d) p_{j,t} = E_t (r_{t+1}^b - r_{t+1}^d) p_{j,t} + \varpi(\omega_{j,t} - \bar{\omega}) n_{j,t}.
\]

Dividing through by \( n_{j,t} \) and re-writing yields:

\[
E_t (r_{t+1}^k - r_{t+1}^b) \phi_t = \varpi(\omega_{j,t} - \bar{\omega}).
\]

Accordingly, given the leverage ratio \( \phi_t \), in the optimum the differential of the expected returns on the individual assets, i.e. \( E_t (r_{t+1}^k - r_{t+1}^b) \), is driven to zero with a speed that is inversely related to the marginal portfolio adjustment costs.

### 4.2.3 Goods-producing firms

The production side of the economy is characterized by four types of firms that are owned by the households: (i) a continuum of perfectly competitive intermediate goods firms indexed by \( i \in [0, 1] \) that produce differentiated goods \( y_{i,t} \), (ii) a continuum of monopolistically competitive retail firms indexed by \( f \in [0, 1] \) that re-package intermediate goods \( y_{i,t} \) into retail goods \( y_{f,t} \), (iii) a continuum of perfectly competitive final goods producers that combine the intermediate goods into a single good \( y_t \), and (iv) a continuum of competitive capital goods producers that repair depreciated capital and build new productive capital.
Final goods producers

A representative final goods firm combines intermediate goods bought from retailers using the technology \( y_t^{(-1)/\epsilon} = \int_0^1 y_{f,t}^{(-1)/\epsilon} df \), where \( \epsilon \) is the elasticity of substitution among intermediate goods. The final goods firm operates in a perfectly competitive market, maximizing profits \( P_t y_t - \int_0^1 P_{f,t} y_{f,t} df \) over input demands \( y_{f,t} \), taking the retail prices \( P_{f,t} \) and the final goods price \( P_t \) as given.

The first-order conditions corresponding to the solution of this problem yield input demand functions, \( y_{f,t} = (P_{f,t}/P_t)^{-\epsilon} y_t \), for all \( f \), and an expression for the aggregate price level, \( P_t^{1-\epsilon} = \int_0^1 P_{f,t}^{1-\epsilon} df \).

Retail firms

Retail firms buy intermediate goods \( y_{i,t} \) at the market price \( P_m \) and re-package those goods into retail goods \( y_{f,t} \) which are sold in a monopolistically competitive market. It takes one unit of intermediate output to make a unit of retail output, i.e. \( y_{f,t} = y_{i,t} \).

The nominal profit of retailer \( f \) in period \( t \) is thus given by \( (P_{f,t} - P_m) y_{f,t} \).

Following Calvo (1983) and Yun (1996), in each period a fraction \( 1 - \psi \) of firms can optimally reset their prices, where \( \psi \) is exogenously given. A firm that can optimally reset its price maximizes the expected sum of discounted profits. The stochastic discount factor for nominal payouts to households is given by \( \beta \Lambda_{t+s}(P_t/P_{t+s}) \), for \( s \geq 0 \). The relevant part of firm \( f \)'s optimization problem is then as follows:

\[
\max_{P_{f,t}} E_t \sum_{s=0}^{\infty} (\beta \psi)^s \Lambda_{t+s}(P_t/P_{t+s}) [P_{f,t} - P_{m,t+s}] y_{f,t+s},
\]

subject to the demand function \( y_{f,t} = (P_{f,t}/P_t)^{-\epsilon} y_t \). By symmetry, all optimizing firms will set the same price \( P_t^* \).

Defining the relative price \( m_t = P_{m,t}/P_t \), the first-order condition is given by

\[
\frac{P_t^*}{P_t} = \frac{\epsilon}{\epsilon - 1} \frac{E_t \sum_{s=0}^{\infty} (\beta \psi)^s \lambda_{t+s} P_{t+s} P_{t+s}^{-\epsilon} m_{t+s} y_{t+s}}{E_t \sum_{s=0}^{\infty} (\beta \psi)^s \lambda_{t+s} P_{t+s}^{-1} P_{t+s}^{1-\epsilon} y_{t+s}}.
\]

Defining further the relative price \( \pi_t^* = P_t^*/P_t \) and the gross inflation rate \( \pi_t = P_t/P_{t-1} \),
the first-order condition can be re-written in recursive form as follows:

\[
\pi_t^* = \frac{\epsilon}{\epsilon - 1} \Xi_{1,t}^t, \quad \Xi_{1,t} = \lambda_t m_t y_t + \beta \psi E_t \pi_{t+1}^{E_t} \Xi_{1,t+1}, \quad \Xi_{2,t} = \lambda_t y_t + \beta \psi E_t \pi_{t+1}^{E_t} \Xi_{2,t+1}.
\]

(4.10)

Finally, by Calvo pricing, the aggregate price level evolves as follows (see Yun, 1996):

\[
1 = (1 - \psi) (\pi_t^*)^{1-\epsilon} + \psi \pi_t^{E_t-1}.
\]

(4.11)

**Intermediate goods producers**

Intermediate goods firms produce differentiated goods that are sold in a perfectly competitive market. Each firm \(i\) has access to the following production technology:

\[
y_{i,t} = a_t (\xi_t k_{i,t-1})^{\alpha} h_{i,t}^{1-\alpha}, \quad \log x_t = \rho_x \log x_{t-1} + \varepsilon_{x,t}, \quad \rho_x \in [0, 1),
\]

for \(x = a, \xi\) and with \(\varepsilon_{x,t} \sim N(0, \sigma_x^2)\). Here, \(a_t\) denotes total factor productivity and \(\xi_t\) denotes the quality of capital. Thus, \(\xi_t k_{i,t-1}\) measures the effective quantity of capital usable for production in period \(t\). The shock \(\xi_t\) is meant to capture economic depreciation or obsolescence of capital and provides a simple source of variation in the quality of capital and thus the value of intermediary assets in the general equilibrium (see Gertler and Karadi, 2011).

Each period, firm \(i\) rents labor services \(h_{i,t}\) at the wage rate \(w_t\) from households and finances its capital acquisition by obtaining funds from financial intermediaries. The timing is as follows. At the end of period \(t\), the firm acquires capital \(k_{i,t}\) for use in production in period \(t+1\). To finance the capital acquisition, the firm issues claims \(s_{k_{i,t}}\) to intermediaries equal to the units of capital acquired, which pay a state-contingent net real return \(r_{t+1}^{k}\) at the beginning of period \(t+1\). The price of each claim is the relative price of a unit of capital \(q_t\). After production in period \(t+1\), the firm sells the effective capital that has depreciated during that period, \((1-\delta)\xi_{t+1} k_{i,t}\), at the price \(q_{t+1}\). Thus, firm \(i\)'s real profits in period \(t\) are given by \(\Pi_{i,t} = m_t a_t (\xi_t k_{i,t-1})^{\alpha} h_{i,t}^{1-\alpha} + q_t (1 - \delta)\xi_t k_{i,t-1} - (1 + r_t^k)q_{t-1} k_{i,t-1} - w_t h_{i,t}\). Taking the relative output price \(m_t\) and the input prices \(q_t, r_t^k\), and \(w_t\) as given, intermediate goods firms maximize \(E_t \sum_{s=0}^{\infty} \beta^s A_{t,t+s} \Pi_{i,t} \).
The first-order conditions are as follows:

\[ h_{i,t} : w_t = (1 - \alpha) m_t y_{i,t} / h_{i,t}, \]

\[ k_{i,t} : E_t \beta \Lambda_{t,t+1} q_t (1 + r^k_{t+1}) = E_t \beta \Lambda_{t,t+1} [\alpha m_{t+1} y_{i,t+1} / k_{i,t} + q_{t+1}(1 - \delta) \xi_{t+1}] . \]

Perfect competition implies that each intermediate goods firm earns zero profits period by period. Accordingly, the firms pay out the ex-post return on capital to financial intermediaries, which is seen by substituting out \( w_t \) in the zero profit condition, i.e. \( \Pi_{i,t} = 0: \)

\[ r^k_t = q_t^{-1} [\alpha m_{t-1} y_{i,t}/k_{i,t-1} + q_t(1 - \delta) \xi_t] - 1. \]

Solving the last expression and the first-order condition for \( h_{i,t} \) for the factor demands yields

\[ h_{i,t} = (1 - \alpha) m_t w_t^{-1} y_{i,t}, \]
\[ k_{i,t-1} = \alpha m_t [q_{t-1}(1 + r^k_t) - q_t(1 - \delta) \xi_t]^{-1} y_{i,t}. \]

Inserting the factor demands into the technology constraint then yields the following expression for the relative intermediate output price:

\[ m_t = \alpha^{-\alpha}(1 - \alpha)^{\alpha-1} a_t^{1-\alpha} [w_t^{1-\alpha} [q_{t-1}(1 + r^k_t) \xi_t^{-1} - q_t(1 - \delta)]^\alpha}. \quad (4.12) \]

### 4.2.4 Capital-producing firms

After production in period \( t \), competitive capital producers purchase the stock of depreciated capital, given by \( (1 - \delta) \xi_t k_{t-1} \), from intermediate goods firms at the relative price \( q_t \). The capital producers combine the depreciated capital with investment goods to produce new productive capital, using an identical capital accumulation technology. The newly produced capital is then sold back to intermediate goods firms and any profits are transferred to households. A representative capital producer’s accumulation technology is given by

\[ k_t = (1 - \delta) \xi_t k_{t-1} + [1 - \Psi(\iota_t)] \iota_t, \quad \Psi(\iota_t) = \frac{\gamma}{2} (\iota_t - 1)^2, \quad \gamma \geq 0, \quad \delta \in [0, 1], \quad (4.13) \]
where \( i_t \) denotes investment expenditures in terms of the final good as a materials input, with relative price unity, and \( \Psi(\cdot) \) are convex investment adjustment costs in \( i_t = i_t/i_{t-1} \). Thus, the capital producer’s real profits in period \( t \) are given by \( q_t k_t - q_t(1 - \delta)\xi_t k_{t-1} - i_t \). The problem of the capital producer is then to solve

\[
\max_{i_t} E_t \sum_{s=0}^{\infty} \beta^s \Lambda_{t,t+s} \{q_{t+s}[1 - \Psi(i_{t+s})] - 1\} i_{t+s},
\]

taking \( q_t \) as given. The first-order condition is as follows:

\[
q_t [1 - \Psi(i_t)] - 1 - qt\Psi'(i_t) + \beta E_t \Lambda_{t,t+1} q_{t+1}i_{t+1}\Psi'(i_{t+1}) = 0,
\]

where \( \Psi'(i_{t+s}) \) denotes the partial derivative of \( \Psi(\cdot) \) with respect to \( i_{t+s} \) for \( s \geq 0 \).

Substituting out the functional terms, the price of capital is seen to satisfy

\[
\frac{1}{q_t} = 1 - \frac{\gamma}{2} \left( \frac{i_t}{i_{t-1}} - 1 \right)^2 - \frac{\gamma i_t}{i_{t-1}} \left( \frac{i_t}{i_{t-1}} - 1 \right) + \beta E_t \Lambda_{t,t+1} q_{t+1}i_{t+1}\Psi'(i_{t+1}) \gamma \left( \frac{i_{t+1}}{i_t} - 1 \right).
\]

(4.14)

### 4.2.5 Fiscal policy

The government conducts purchases of the final good and financial sector policies. Government purchases \( g_t \) consist of a stochastic part \( \tilde{g}_t \) plus a possible response to shocks \( \xi_t \):

\[
g_t = \tilde{g}_t + \varsigma(\xi_{t-l} - \xi), \quad \varsigma \leq 0, \quad l \geq 0,
\]

where \( \tilde{g}_t \) follows an autoregressive process in logs, \( \log(\tilde{g}_t/\bar{g}) = \rho_g \log(\tilde{g}_{t-1}/\bar{g}) + \varepsilon_{u,t}^g + \varepsilon_{a,t}^g \), with \( \varepsilon_{x,t}^g \sim N(0,\sigma_{x,t}^2) \) for \( x = u, a \), \( \rho_g \in [0,1) \), and \( \bar{g} > 0 \). Below, we study the effects of surprise changes in spending due to the unanticipated shock \( \varepsilon_{u,t}^g \) and spending changes that are pre-announced one year in advance due to the news shock \( \varepsilon_{a,t}^g \). The parameter \( \varsigma \) determines the spending response to shocks to the quality of capital. Below, this shock serves as the initiating disturbance leading to a financial crisis. If \( \varsigma < 0 \), spending increases during the crisis above its steady state value. If \( \varsigma = 0 \), there is no government intervention. Through the parameter \( l \), the response occurs contemporaneously (\( l = 0 \)) or with some lag (\( l > 0 \)). Although it may seem
less appealing from a practical point of view than e.g. an endogenous output feedback, an exogenous feedback of this type makes the policy experiments conducted below comparable by excluding second-round effects onto government interventions.

We also allow for different kinds of government interventions in the financial sector. In particular, we assume that the government is willing to provide funds $n_{g,t}$ to financial intermediaries according to a rule that is symmetric to the spending rule:

$$n_{g,t} = \kappa (\xi_{t-1} - \xi), \quad \kappa \leq 0, \quad l \geq 0.$$  

According to this rule, if $\kappa < 0$, the government provides funds in the face of shocks to the quality of capital, which can again occur contemporaneously or with some lag as determined by $l$. If $\kappa = 0$, there are no government interventions in the financial sector. In addition, we allow for the possibility that the funds provided in this way are repaid by the intermediaries, where the period-$t$ repayment $\hat{n}_{g,t}$ is specified as follows:

$$\hat{n}_{g,t} = \vartheta n_{g,t-1} - e, \quad \vartheta \geq 0, \quad e \geq 1.$$  

Hence, the size of intermediary repayments relative to the funds provided by the government is determined by the penalty factor $\vartheta$: if $\vartheta = 0$, the government makes a transfer or “gift” to intermediaries, $\vartheta = 1$ nests the case of a zero-interest loan, and if $\vartheta > 1$, the funds need to be repaid at some positive (penalty) interest rate. Furthermore, any repayments occur with some delay as determined by the parameter $e$.

Let $b_{t-1} (b_t)$ denote the stock of government debt at the beginning (at the end) of period $t$. We assume that, perhaps hypothetically, the government can raise lump-sum taxes from households according to the rule

$$\tau_t = \bar{\tau} + \kappa_b (b_{t-1} - b) + \kappa_g (g_t - g) + \kappa_n n_{g,t}, \quad \kappa_b > 0, \quad \kappa_g, \kappa_n \in [0, 1], \quad \bar{\tau} > 0.$$  

As $\kappa_b > 0$, following Bohn (1998), a tax rule of this type ensures fiscal solvency for any finite initial level of debt. In addition, we would like to have a benchmark against

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5Similar types of policies have been implemented during the recent crisis e.g. in the Netherlands, where penalty interest rates up to 50 percent were charged on government loans to financial institutions that were to be repaid after about three years.
which we can judge the effects of intermediary deficit financing. We therefore introduce
the parameters $\kappa_g$ and $\kappa_n$ and the respective components. For $\kappa_g = 0$ ($\kappa_n = 0$), goods
purchases (financial sector policies) are fully financed by deficits. For $\kappa_g = 1$ and
$\kappa_n = 1$, fiscal policy is entirely financed by lump-sum taxes on households. It should
be noted that the introduction of small distortionary taxes for debt repayment are
unlikely to change our main qualitative conclusions.

The period-by-period government budget constraint is then given by

$$b_t + \tau_t + \tilde{n}_{g,t} = g_t + n_{g,t} + (1 + r^b_t)b_{t-1}. \quad (4.15)$$

### 4.2.6 Monetary policy

To close the model, we assume that the monetary authority sets the risk-free nominal
interest rate on deposits $r^n_t$ to stabilize inflation and output according to a Taylor rule
of the form

$$r^n_t = (1 - \rho_r) \left[ r^n + \kappa_\pi (\pi_t - \bar{\pi}) + \kappa_y \log(y_t/y_{t-1}) \right] + \rho_r r^n_{t-1} + \varepsilon_{r,t}, \quad \kappa_y \geq 0, \quad \kappa_\pi > 1,$$

with $\rho_r \in [0, 1)$ and $\varepsilon_{r,t} \sim N(0, \sigma^2_r)$. The parameter $\bar{\pi} \geq 1$ stands for the inflation
target. The strength of the monetary authority’s reaction to fluctuations of inflation
and output is determined by the parameters $\kappa_\pi$ and $\kappa_y$, where we have imposed Taylor’s
(1993) principle as $\kappa_\pi > 1$. We also allow for an interest rate smoothing component
in the Taylor rule, where the strength of interest rate smoothing is controlled by the
parameter $\rho_r$. The following Fisher relation then defines the ex-post gross real interest
rate on deposits:

$$1 + r^d_t = (1 + r^n_{t-1}) \pi_t^{-1}. \quad (4.16)$$

Notice that the model emphasizes the direct links of e.g. central bank lending rates
to intermediary funding rates by the choice of the deposit rate as the instrument for
monetary policy. The interest rate on government bonds is however endogenously
determined in the general equilibrium.

---

6The specification of the Taylor rule uses the log deviation of current output from last period’s output, to approximate the output gap that appears in Taylor’s (1993) original version, following common specifications in empirical macroeconomic models (e.g. Christoffel, Coenen, and Warne, 2008).
4.2.7 Aggregation and market clearing

Financial variables

Given the overall asset size, $p_{j,t} = \phi_t n_{j,t}$, and the asset structure of the balance sheets of individual financial intermediaries, $q_t s^k_{j,t} = \omega_{j,t} \phi_t n_{j,t}$ and $s^b_{j,t} = (1 - \omega_{j,t}) \phi_t n_{j,t}$, the evolution of intermediary $j$’s net worth can be re-written as follows:

$$n_{j,t+1} = [(r^p_{t+1} - r^d_{t+1}) \phi_t + 1 + r^d_{t+1} - \Omega(\omega_{j,t})] n_{j,t}.$$  

We therefore also have the terms

$$x_{t,t+1} = n_{j,t+1}/n_{j,t} = (r^p_{t+1} - r^d_{t+1}) \phi_t + 1 + r^d_{t+1} - \Omega(\omega_{j,t}),$$

$$z_{t,t+1} = p_{j,t+1}/p_{j,t} = (\phi_{t+1}/\phi_t) (n_{j,t+1}/n_{j,t}) = (\phi_{t+1}/\phi_t) x_{t,t+1},$$

$$f_{t,t+1} = d_{j,t+1}/d_{j,t} = (\phi_{t+1} - 1)/(\phi_t - 1) (n_{j,t+1}/n_{j,t}) = (\phi_{t+1} - 1)/(\phi_t - 1) x_{t,t+1},$$

The portfolio problem of intermediary $j$ further implies that the individual portfolio weights are given by

$$\omega_{j,t} = \omega - E_t_r (r^k_{t+1} - r^b_{t+1}) \phi_t + \bar{\omega}.$$  

Substituting out the latter as well as $v_t, \eta_t$, and $\phi_t$ in the above terms, it follows that none of the components of $\phi_t$ depend on individual factors. Thus, we also have that $\omega_{j,t} = \omega_t$ for all $j$.

The aggregate asset demands $s^k_t = \int_0^1 s^k_{j,t} dj$ and $s^b_t = \int_0^1 s^b_{j,t} dj$ then follow as

$$q_t s^k_t = \omega_t \phi_t n_t, \quad s^b_t = (1 - \omega_t) \phi_t n_t,$$

where $n_t = \int_0^1 n_{j,t}$ denotes aggregate net worth. Aggregate net worth $n_t$ is the sum of total net worth of financial intermediaries that continue operating $n_{c,t}$, total net worth of newly entering intermediaries $n_{e,t}$, and net transfers by the government, $n_{g,t} - \tilde{n}_{g,t}$.

Total net worth of continuing intermediaries is given by $n_{c,t} = \theta [ (r^p_t - r^d_t) \phi_{t-1} + 1 + r^d_t - \Omega(\omega_{t-1})] n_{t-1}$. To obtain an expression for $n_{e,t}$, it is assumed that new bankers receive a start-up transfer from households equal to a fraction $\chi/(1 - \theta)$ of aggregate net worth at the end of period $t - 1$, which is equal to $(1 - \theta)n_{t-1}$. Thus, $n_{e,t} = \chi n_{t-1}$. 


Accordingly, we have

\[ n_t = \{\theta[(r_t^p - r_t^d)\phi_{t-1} + 1 + r_t^d - (\varpi/2)(\omega_{t-1} - \bar{\omega})^2] + \chi\}n_{t-1} + n_{g,t} - \tilde{n}_{g,t}, \quad (4.18) \]

Further, aggregate securities issued by intermediate goods firms to financial intermediaries satisfy

\[ q_t \int_0^1 s_{i,t}^k di = q_t \int_0^1 s_{j,t}^k dj = q_t \int_0^1 k_{i,t} di, \quad \text{or, using the market clearing conditions} \quad s_t^k = \int_0^1 s_{i,t}^k di = \int_0^1 s_{j,t}^k dj \quad \text{and} \quad k_t = \int_0^1 k_{i,t} di: \]

\[ s_t^k = k_t. \quad (4.19) \]

Similarly, aggregate bonds issued by the government to financial intermediaries satisfy

\[ s_t^b = b_t. \quad (4.20) \]

The aggregate asset portfolio follows by integrating over individual portfolios:

\[ p_t = \int_0^1 p_{j,t} dj = q_t \int_0^1 s_{j,t}^k dj + \int_0^1 s_{j,t}^k dj = q_t s_t^k + s_t^b. \quad (4.21) \]

Aggregate deposits follow by integrating over individual balance sheets:

\[ d_t = \int_0^1 d_{j,t} dj = \int_0^1 p_{j,t} dj - \int_0^1 n_{j,t} dj = p_t - n_t. \quad (4.22) \]

**Factor demands**

Demand by final goods producers for each retail good is \( y_{f,t} = y_{i,t} = y_t(P_{f,t}/P_t)^{-\epsilon} \), for all \( f \) and all \( i \). With \( y_{i,t} = y_{f,t} \), the factor demands by firm \( i \) are given by

\[ h_{i,t} = (1 - \alpha)m_t w_t^{-1} y_{f,t}, \quad k_{i,t-1} = \alpha m_t [q_{t-1}(1 + r_t^k) - q_t(1 - \delta)\xi_t]^{-1} y_{f,t}. \]

The aggregate factor demands follow by the market clearing conditions \( \int_0^1 h_{i,t} di = h_t \)

and \( \int_0^1 k_{i,t-1} di = k_{t-1} \):

\[ h_t = (1 - \alpha)m_t w_t^{-1} y_t \Delta_t, \]

\[ k_{t-1} = \alpha m_t [q_{t-1}(1 + r_t^k) - q_t(1 - \delta)\xi_t]^{-1} y_t \Delta_t, \]
where $\Delta_t = \int_0^1 (P_{f,t}/P_t)^{-\epsilon} df$ is a price dispersion term with the recursive form

$$\Delta_t = (1 - \psi) (\pi_t^*)^{-\epsilon} + \psi \pi_t^{\epsilon} \Delta_{t-1},$$  \hspace{1cm} (4.23)

see Yun (1996). Hence, the aggregate capital-labor ratio follows as

$$k_{t-1}/h_t = \alpha (1 - \alpha)^{-1} w_t[q_{t-1}(1 + r_t^b) - q_t(1 - \delta)\xi_t]^{-1} = k_{i,t-1}/h_{i,t}. \hspace{1cm} (4.24)$$

### Aggregate supply

Integrating $y_{i,t} = a_t(\xi_t k_{i,t-1})^\alpha h_{i,t}^{1-\alpha}$ over $i$, it follows that

$$\int_0^1 a_t(\xi_t k_{i,t-1})^\alpha h_{i,t}^{1-\alpha} di = a_t \xi_t \left( \frac{k_{t-1}}{h_t} \right)^\alpha \int_0^1 h_{i,t} di = a_t(\xi_t k_{t-1})^\alpha h_t^{1-\alpha}.$$

Integrating $y_{f,t} = y_t(P_{f,t}/P_t)^{-\epsilon}$ over $f$ then yields output of the final good:

$$y_t \Delta_t = a_t(\xi_t k_{t-1})^\alpha h_t^{1-\alpha}. \hspace{1cm} (4.25)$$

### Goods market clearing

Goods market clearing further requires that aggregate demand equals aggregate supply:

$$c_t + i_t + g_t = y_t. \hspace{1cm} (4.26)$$

#### 4.2.8 Equilibrium

The rational expectations equilibrium of this model is then the set of sequences $\{c_t, h_t, w_t, i_t, k_t, q_t, y_t, m_t, \pi_t, \pi_t^*, \Xi_t, \Xi_{2,t}, \Delta_t, \tau_t, \rho_t, \lambda_t, \omega_t, \xi_t, q_t, \phi_t, \eta_t, n_t, s_t^b, s_t^k, p_t, d_t, b_t\}_{t=0}^{\infty}$ and shadow prices $\{\lambda_t\}_{t=0}^{\infty}$, such that for given initial prices and initial values, a fiscal policy $\{g_t, n_{g,t}, \bar{n}_{g,t}, \tau_t\}_{t=0}^{\infty}$, a monetary policy $\{\rho_t\}_{t=0}^{\infty}$, and sequences of shocks $\{a_t, \xi_t\}_{t=0}^{\infty}$, conditions (4.1)-(4.26), dropping the $j$ subscripts for individual intermediaries where appropriate, and the transversality conditions are satisfied.

This closes the description of the model. The model is solved by a first-order perturbation around the non-stochastic steady state which is derived in Appendix 4.A.
4.3 Model analysis

4.3.1 Calibration

Table 4.1 lists the choice of parameters for the baseline version of the model. The calibration mostly follows Gertler and Karadi (2011). This concerns the subjective discount factor \( \beta \), the degree of habit formation \( \nu \), the Frisch elasticity of labor supply \( \varphi^{-1} \), the elasticity of substitution among intermediate goods \( \epsilon \), the Calvo probability of keeping prices fixed \( \psi \), the effective capital share in production \( \alpha \), and the investment adjustment cost parameter \( \gamma \). The respective parameter values are estimates by Primiceri, Schaumburg, and Tambalotti (2006). The parameters in the monetary policy rule are set to conventional values. In addition, we take a conservative stance on the parameters that are specific to our model. In particular, we use a small value for the portfolio adjustment cost parameter \( \omega \) to limit the impact of the adjustment costs on the dynamics to a minimum (cf. Footnote 8). The value of the debt feedback on taxes \( \kappa_b \) is chosen to have stability conditions satisfied in both the version of the model with intermediaries and the version without intermediaries (see Appendix 4.B).

Following again Gertler and Karadi (2011), the steady state leverage ratio \( \phi \) is set to four to roughly match aggregate U.S. financial data. The steady state credit spread \( \Gamma \) is set to one hundred basis points to match the pre-2007 spreads of bank lending rates to risk-free bonds. The average survival rate of bankers \( \Theta = 1/(1 - \theta) \) is set to sixteen quarters (thus smaller than in Gertler and Karadi, 2011) by calibrating the survival probability \( \theta \), to make sure that the proportional transfer to entering bankers \( \chi \) is positive (see Appendix 4.A). To roughly match U.S. macroeconomic data, the steady state ratios of investment and government spending over GDP \( i/y \) and \( g/y \) are set to 20 percent, the latter by calibrating \( \delta \), and the ratio \( b/y \) is set to 2.4 which implies an annual debt-to-GDP ratio of 60 percent.

4.3.2 Surprise spending shock

We begin our discussion of results by an examination of the dynamics due to a surprise spending shock. Figure 4.1 shows the responses of selected variables to an unanticipated
Table 4.1: Model parameters

<table>
<thead>
<tr>
<th>Parameter</th>
<th>Value</th>
<th>Definition</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Households</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$\beta$</td>
<td>0.990</td>
<td>Subjective discount factor</td>
</tr>
<tr>
<td>$\nu$</td>
<td>0.815</td>
<td>Degree of habit formation</td>
</tr>
<tr>
<td>$\varphi$</td>
<td>0.276</td>
<td>Inverse Frisch elasticity of labor supply</td>
</tr>
<tr>
<td><strong>Financial intermediaries</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$\lambda$</td>
<td>0.226</td>
<td>Fraction of assets that can be diverted</td>
</tr>
<tr>
<td>$\theta$</td>
<td>0.938</td>
<td>Survival probability of bankers</td>
</tr>
<tr>
<td>$\chi$</td>
<td>0.016</td>
<td>Proportional transfer to entering bankers</td>
</tr>
<tr>
<td>$\zeta$</td>
<td>0.001</td>
<td>Portfolio adjustment cost parameter</td>
</tr>
<tr>
<td><strong>Goods-producing firms</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$\epsilon$</td>
<td>4.176</td>
<td>Elasticity of substitution</td>
</tr>
<tr>
<td>$\psi$</td>
<td>0.779</td>
<td>Calvo probability of keeping prices fixed</td>
</tr>
<tr>
<td>$\alpha$</td>
<td>0.330</td>
<td>Share of effective capital in production</td>
</tr>
<tr>
<td><strong>Capital-producing firms</strong></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$\delta$</td>
<td>0.079</td>
<td>Depreciation rate of effective capital</td>
</tr>
<tr>
<td>$\gamma$</td>
<td>1.728</td>
<td>Investment adjustment cost parameter</td>
</tr>
<tr>
<td><strong>Policy</strong>$^a$</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$\kappa_b$</td>
<td>0.020</td>
<td>Government debt feedback on taxes</td>
</tr>
<tr>
<td>$\rho_r$</td>
<td>0.800</td>
<td>Interest rate smoothing parameter</td>
</tr>
<tr>
<td>$\kappa_{\pi}$</td>
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<td>Inflation feedback on nominal interest rate</td>
</tr>
<tr>
<td>$\kappa_y$</td>
<td>0.125</td>
<td>Output feedback on nominal interest rate</td>
</tr>
</tbody>
</table>

An increase in government spending (goods purchases) that is normalized to 1% of GDP on impact and that is persistent with autocorrelation coefficient $\rho_g = 0.8$. We consider the case of full intermediary financing and, as a benchmark, the case of full household financing through lump-sum taxes. The figure also shows the impulse responses from the version of the model without financial intermediaries, as a reference case. This version of the model is described in Appendix 4.B.

According to the model without intermediaries, the spending expansion raises output by more than one percent on impact, since investment increases initially while the fall in consumption that is caused by consumption smoothing in the face of higher future taxes is subdued initially through the presence of habit formation. However, in the model with intermediaries and with intermediary financing of the spending expansion, the output response is smaller than one percent on impact and it is also significantly
Figure 4.1: Impulse responses to a surprise spending shock

Note. Unexpected increase in government spending in quarter 0 (innovation $\varepsilon_{g,t}$) by 1% of GDP relative to its steady state value.

less persistent as output decreases over the medium term.

Underlying those effects are the funding pressures that are put on financial intermediaries by a deficit-financed fiscal expansion. The fiscal expansion raises both expected interest rates through the associated tightening of intermediary balance sheet constraints and intermediary balance sheet adjustments. The precise mechanism through which this occurs is explained in Section 4.3.3 below. As a consequence of the rise in borrowing costs, the demand for capital by intermediate goods firms and thus investment by capital producers is crowded out. The fall in investment is amplified by the
financial accelerator mechanism described in Gertler and Karadi (2011) that is due to procyclical variation in intermediary balance sheets: falling investment leads to a falling price of capital, which lowers intermediary net worth and thus further tightens intermediary constraints, which further raises borrowing costs such that investment falls by more, further decreasing asset prices, and so forth. These effects feed through the whole economy as falling wages distract household labor supply and as the associated worsening of budgetary conditions depresses consumption.

It is interesting to see that the spending expansion would be more effective in our model when it would be financed by households compared to the model without intermediaries, where households hold all government bonds and claims issued by intermediate goods firms. Under household financing, the intermediary balance sheet mechanism makes the spending expansion comparably more effective as the build-up of investment (which occurs under household financing just like in the model without intermediaries) raises the price of capital over time and thus eases intermediary balance sheet constraints. As a consequence, an analysis of stimulus policies in a similar model as in Gertler and Karadi (2011), without considering intermediary financing of government deficits, is likely to lead to the conclusion that the benefits of such policies are enhanced by the relevant financial frictions.

4.3.3 Pre-announced spending shock

The intermediary balance sheet adjustments in interaction with balance sheet constraints that are at the heart of our model become clear when we look at the effects of a spending increase that is pre-announced one year in advance. This experiment allows to distinguish the relevant expectational effects from other effects.

Thus, Figure 4.2 shows the response of the economy to news in quarter 0 that spending is going to increase by 1% of GDP in quarter 4. The results show that the spending expansion would have almost no effect on output under household financing after the news and before the implementation, similarly as in the model without intermediaries. With intermediary financing, however, output falls in the first year due to

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7In that model, borrowing rates co-move through arbitrage behavior by households and the central bank sets the interest rate on bonds. The model is observationally equivalent to one where financial intermediaries would take deposits from households but face no leverage constraints.
an immediate fall in investment after the announcement of the spending expansion.

The underlying mechanism is revealed by a closer look at the impulse responses of a second set of variables to the same shock which are shown in Figure 4.3. A first effect works through a tightening of intermediary balance sheet constraints. The expected future increase in government primary deficits due to the upcoming spending expansion implies higher expected growth rates of bonds and total assets. The intermediaries have an incentive to accumulate assets due to a rise in the expected discounted marginal gain of assets. This incentive, however, tightens leverage constraints as indicated by a
strong rise in the associated Lagrangian multiplier. The latter restricts intermediary asset demand and raises the costs of credit to both the government and intermediate goods firms during the announcement period (i.e. it raises credit spreads).

A second effect works through intermediary portfolio adjustments. Rising spreads distract investment, which lowers the price of capital. Everything else equal, the fall in the price of capital enhances the rise in the expected return on capital. Intermediaries thus shift their portfolios into assets with higher expected returns, i.e. claims on intermediate goods firms. The associated fall in the demand for bonds reduces the implicit bond price and raises the ex-ante nominal interest rate on bonds. The expected

Note. See Figure 4.2.
real rate on bonds increases, which adds to the rise in the expected overall portfolio return. The increase in the expected portfolio return further enhances incentives to accumulate assets and thus reinforces the first effect discussed above.\footnote{The portfolio shift into claims tends to dampen the rise in the expected return on capital. With higher portfolio adjustment costs, this dampening effect becomes weaker and the crowding-out effect stronger. By allowing for low adjustment costs, our results thus rather fall on the conservative side.}

Notice that the fall in the demand for bonds is only consistent with an equilibrium if the ex-post real interest rate on bonds falls given a fixed initial supply of bonds by the government. Hence, a planned fiscal expansion in this environment might give an impression of further fiscal space due to low interest rates when, in fact, there is none. Once the expansion takes place output shoots up to give the arguably erroneous impression that it is actually effective, albeit only for a short time.

### 4.3.4 Financial crisis and policy responses

We now analyze the effects of alternative government policy interventions during a simulated financial crisis. The type of crisis is the same as in Gertler and Karadi (2011), the initiating disturbance being an unexpected decline in the quality of capital $\xi_t$ by five percent on impact with autocorrelation coefficient $\rho_\xi = 0.66$.

This experiment creates a similar crisis of the type, the magnitude, and the duration of the recent crisis. Other initiating shocks are conceivable, but the specific type of shock is irrelevant for the qualitative implications. Following Gertler and Karadi (2011), we assume that the monetary authority reduces its tendency to smooth interest rates in the face of the shock, to capture the notion that monetary policy tends to act more aggressively during a financial crisis.\footnote{The smoothing parameter $\rho_r$ in the Taylor rule is reduced by half but not more so that non-negativity constraints on nominal interest rates are satisfied.} As seen in the following figures, absent any government intervention, the deterioration in intermediary asset quality produces a sharp recession with a peak output decline of more than five percent, as intermediary net worth drops and credit tightens, leading to a sharp rise of the credit spread. As investment eventually picks up when financial conditions have calmed down to rebuild the destroyed capital stock, it takes more than four years for output to recover.

We consider the following government policies: (i) immediate deficit-financed spend-
Figure 4.4: Crisis policy I – immediate spending stimulus

Note. Initiating shock is unexpected decline in the quality of capital by five percent relative to its steady state value in quarter 0; autocorrelation coefficient $\rho_\xi = 0.66$.

...ing stimulus; (ii) delayed stimulus; (iii) immediate transfers to intermediaries; (iv) delayed transfers; (v) zero-interest loans with delayed repayment; (vi) loans at penalty interest rates; (vii) partly and fully tax-financed spending stimulus. This set of policies is sufficient to explain the relevant implications of the model, but it also suitably captures the main fiscal measures applied during the recent crisis. The policies are either financed by issuing bonds to intermediaries or by raising lump-sum taxes directly from households without resorting to the intermediaries, as a benchmark.

To make the results comparable, the different policy measures are scaled to have the same size relative to GDP on impact by adjusting the feedback parameters $\varsigma$ and $\kappa$ accordingly.
Financial stress, Gov. Policies, and the Consequences of Deficit Financing

**Fiscal stimulus**

Figure 4.4 illustrates the effects of a countercyclical, persistent spending stimulus of two percent of GDP that occurs immediately when the shock hits. The results show that the deficit-financed demand stimulus is able to dampen the initial decline in output by more than one percentage point relative to the no intervention case. After about one year, the fall in output turns however slightly stronger than without government intervention and it remains more negative afterwards. This effect is due to an enhanced fall in investment associated with an additional rise in the credit spread that is caused by a further tightening of leverage constraints due to the increase in government borrowing. The fall in investment and an enhanced decline in consumption due to a higher tax load offset the output gain from the additional government spending.

What happens if the spending stimulus occurs relatively late during the recession? In Figure 4.5 the policy is implemented with a delay of one year after the initial shock. The deficit-financed stimulus seems counterproductive in this case as the fall in output is amplified. The peak decline in output now reaches almost six percent. The reason is that credit tightens immediately in the face of the upcoming spending expansion, as discussed in Section 4.3.3. A similar fall in investment as under an early fiscal expansion thus takes place, but the actual stimulus arrives later, such that the initial decline in output is further amplified compared to the no intervention case.

Notice that in both of the above cases the stimuli would be more effective under the benchmark of household financing than under deficit financing. The reasoning goes in line with the discussion in Sections 4.3.2 and 4.3.3, the latter because the upcoming stimulus is anticipated by the agents as soon as the capital quality shock hits.\(^\text{11}\)

**Financial sector support**

We analyze next the effects of financial sector support measures, in the form of transfers to intermediaries with possible repayment. The responses of the credit spread \(E_t[r^k_{t+1} - r^d_{t+1}]\) and output under pure transfers, immediate or delayed by one year, and zero-

\(^{11}\)A household-financed stimulus would also lead to a small decrease in output during the initial quarters, as seen in Figure 4.4. This result can be adhered to the lower interest smoothing parameter in the crisis experiment and the associated valuation effects on intermediary liabilities.
interest loans with delayed repayment after one year are shown in Figure 4.6. The charts in the first two rows show that both immediate pure transfers but also delayed pure transfers, due to the forward-looking character of the intermediary constraints, are able to moderate the recession. The positive effect on output is stronger under household financing, but also under deficit financing the policy tends to be beneficial. In both cases, the transfers dampen the rise in the spread by raising intermediary net worth. The third row considers the case of zero-interest loans that are repaid after four quarters. According to the results, such loans are not effective in dampening the crisis under deficit financing. In the benchmark case of household financing, however, the
Figure 4.6: Crisis policies III to V – immediate transfers to intermediaries, delayed transfers, and zero interest loans with delayed repayment

Notes. III: immediate transfers; IV: delayed transfers; V: zero interest loans; see Figure 4.4.

loans would be effective in reducing the output loss.

We can also go a step further and look at loans that need to be repaid at penalty interest rates. Figure 4.7 shows the dynamics due to the capital quality shock when the loans need to be repaid after sixteen quarters, at zero interest or at one hundred and two hundred percent penalty rates.\textsuperscript{12} The figure also shows the dynamics without government intervention (dashed line) and with transfers to intermediaries (thickest solid line). The figure shows that the downturn is dampened during the initial quarters,

\textsuperscript{12}The delay of sixteen quarters is chosen because at that point the crisis is arguably over, thus motivating repayment of temporary support measures.
Figure 4.7: Crisis policy VI – loans at penalty interest rates

Note. See Figure 4.4.

the more the higher the repayment, but towards the time of the repayment the recession is prolonged in a double-dip fashion, the second (first) dip being larger (smaller) the larger the repayment requirements. In fact, the credit spread shows spikes around the time when the repayments are due in the four different cases.

It seems interesting that the initial recession is dampened more strongly the higher the repayment factors. This effect is due to an initial easing of credit conditions. The main underlying reason is that capital producing firms anticipate the future tightening of credit conditions and the associated future fall in the price of capital. The capital producers therefore increase their initial investment, given relatively higher resale
Figure 4.8: Output gains from penalty-rate loans

Notes. Crisis experiment as in Figure 4.4; 2% of GDP initial loans with delayed repayment at penalty rates; penalty factor \( x \): \( 100 \times (x - 1) \) percent penalty interest rate.

prices, which leads to an accelerated rise in the price of capital that eases balance sheet constraints on financial intermediaries and thus tends to dampen the initial crisis.

Output gains

Can the stabilizing effects during the crisis potentially suffice to generate overall output stabilization gains from penalty loans? Figure 4.8 plots measures of output gains against the penalty factor in loan repayment, which occurs after sixteen quarters. The four charts show the impact responses, the minimum responses, and the undiscounted and discounted cumulative responses of GDP under both deficit-financed loans (dashed
Figure 4.9: Output gains from stimuli and transfers under mixed financing

Notes. Crisis experiment as in Figure 4.4; 2% of GDP initial spending stimuli; degree of tax financing \( x \): 100 \( \times \) \( x \) percent household financing, 100 \( \times \) \((1 - x)\) percent intermediary financing.

lines) and household-financed loans (dashed-dotted lines) relative to the case of no intervention (solid lines).\(^\text{13}\) The loans are again scaled to two percent of GDP.

According to the impact and minimum responses, under both deficit financing and household financing, loan provision is found to be an effective means to dampen the crisis recession for all values of the penalty factor considered. The reason is that some of the output loss is instead shifted towards later periods due to (the anticipation of) rising

\[^{13}\text{Denote as } \tilde{y}_k \text{ the percentage deviation of output from its steady state value at horizon } k = 0, 1, 2, \ldots, T \text{ with the government intervention. The measures are calculated as follows: impact responses } \tilde{y}_0; \text{ minimum responses } \min_k \tilde{y}_k; \text{ undiscounted cumulative responses } \sum_{k=0}^{T} \tilde{y}_k; \text{ discounted cumulative responses } \sum_{k=0}^{T} \beta^k \tilde{y}_k, \text{ where } \beta \text{ is the household subjective discount factor. We set } T = 1000.\]
credit spreads and thus falling investment at the time of the repayment, as indicated by the decreasing undiscounted cumulative gains (see Section 4.3.4). However, most remarkably, under household financing the prediction would be that even very large penalty factors can still bring overall cumulative stabilization gains. Under deficit financing, on the other hand, there are no overall stabilization gains according to both cumulative measures considered.

Hence, a straightforward analysis of financial sector policies in a similar model as in Gertler and Karadi (2011), without intermediary financing of government deficits, could lead to a rather odd conclusion: overall output stabilization gains are possible when temporary support measures are repaid after some time at huge penalty rates. When considering deficit financing by intermediaries, however, the relatively small overall gains from deficit-financed policies should lead to more cautious predictions.

As a final step, we examine stabilization gains from demand stimulus as well as transfers to intermediaries depending on the degree of deficit financing as determined by $\kappa_g$ and $\kappa_n$. This final experiment serves to compare whether there is some critical point at which the benefits from these policies surpass the costs due to the tightening of intermediary constraints, in view of the core question of this chapter. According to the results in Figure 4.9, both measures are least effective under full deficit financing. Transfers, hypothetically perhaps, can bring stabilization gains even under full deficit financing. However, already for moderate degrees of household financing above 20 percent, the stimulus is also able to dampen the recession (cf. the minimum responses) and moderate the overall output loss (cf. the cumulative responses). This result again emphasizes the importance of taking the precise financing mode of fiscal policy into account when deciding on policy measures in a situation of financial stress.

### 4.4 Related empirical evidence

This section provides a brief review of the empirical evidence that is linked to our study. The related evidence can be grouped into results from fiscal VAR studies on the

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14 The kink in the upper right chart is due to the fact that at some point, for relatively high penalty factors, the output drop at the time of the repayment turns larger than the minimum response during the crisis.
effects of government spending or goods purchases, in particular on private investment, findings of cross-sectional studies that investigate the impact of financial sector policies and fiscal stimulus during financial crises, and results from empirical studies looking at the effects of fiscal finances on interest rates.

Among the first group of studies, both structural VAR (SVAR) methods and event-study approaches point towards negative effects of government spending on private investment. On the SVAR side, Blanchard and Perotti (2002) find that investment is consistently crowded out by government spending shocks in the U.S. over the period 1960Q1-1997Q4, with a peak decline of up to one percent due to a 1% of GDP spending increase. Using a yearly panel VAR on 18 OECD countries over the period 1960-1996, Alesina, Ardagna, Perotti, and Schiantarelli (2002) also find a sizable negative effect of public spending (particularly public wages) on investment, a one percentage point increase in the primary spending-to-GDP ratio leading to a fall in the investment-to-GDP ratio of 0.15 percentage points on impact and to a cumulative fall of 0.74 percentage points after five years. On the side of event-studies, identifying spending shocks based on war dates and professional forecasts, Ramey (2011b) finds that after a positive defense news shock in the U.S. both non-residential and residential investment fall significantly, with peak effects of up to -1 percent (non-residential investment) and -1.5 percent (residential investment). Shocks identified based on professional forecast errors over the period 1969-2008 indicate even stronger falls of -1.5 percent and -3.5 percent, respectively, as well as a medium-term decline in output.

Our model predicts stronger crowding-out effects of spending-based fiscal expansions on investment than most of the above VAR studies (see Figures 4.1 and 4.2).15 The qualitative predictions are however similar. The quantitative differences also do not come as a big surprise as the model mainly describes business cycles in times of financial stress, whereas the above studies look instead at sample averages. Related to this fact, Baldacci, Gupta, and Mulas-Granados (2008) estimate the effects of fiscal policy interventions during 118 episodes of banking crises in a cross-section of developed and emerging countries. In line with the results of Section 4.3.4, they find that

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15The fall in investment in Figure 4.1 implies a 0.48 percentage points decline in the investment-to-GDP ratio on impact and a cumulative fall of more than three percentage points after five years.
financial sector policies can shorten such crises whereas fiscal stimulus going along with such policies can have stabilizing effects, but the latter does not hold for countries where fiscal policy is subject to funding constraints.\footnote{A similar conclusion is reached by Ilzetzki, Mendoza, and Végh (2010); based on a quarterly dataset for 20 developed and 24 developing countries over the period 1960-2007, they find that during episodes where government debt was higher than 60% of GDP, spending multipliers are not statistically different from zero on impact and negative (and statistically significant) in the long run.}

Finally, a recent study on the effects of fiscal finances on interest rates has been conducted by Laubach (2009), who estimates the effects of U.S. government debt and deficits on Treasury yields, isolating those effects from other factors affecting interest rates (e.g. due to countercyclical monetary policy and automatic fiscal stabilizers) by focusing on the relation between long-horizon expectations of both interest rates and fiscal variables. According to Laubach, the idea is that measures of expectations hold out the prospect of uncovering causal effects from fiscal variables to interest rates. Laubach concludes that the effects of fiscal variables on interest rates are statistically significant and economically relevant; in particular, for the period 1976-2006 an increase in the projected deficit-to-GDP ratio by one percentage point raises forward rates five and more years into the future by about 25 basis points. This result corroborates the findings shown in Figures 4.1 and 4.2 from a qualitative perspective.\footnote{See also Canzoneri, Cumby, and Diba (2002), Cohen and Garnier (1991), Elmendorf (1993), Wachtel and Young (1987), as reviewed in Laubach (2009). Most of these studies also find statistically and economically significant effects of deficits on interest rates.}

### 4.5 Conclusion

After a wave of calls for fiscal stimulus to lift distressed economies out of recession in the context of the recent crisis, when economic growth finally picked up after the implementation of large fiscal packages, the policy efforts where deemed effective at first.\footnote{See, for instance, Romer (2009), Romer and Bernstein (2009); see also CBO (2009), IMF (2009a), and OECD (2009).} However, despite reduced credit spreads and diminishing job losses, the recovery turned out to be less robust than originally hoped, particularly in the U.S. where the largest package was adopted in absolute size (see Mankiw, 2009). Dissatisfaction with the impact of the implemented policies spread and many observers asked why the stimulus was not more effective (see e.g. Adams and Gangnes, 2011).
Our model is able to provide an answer to this question. Key to the answer is the fact that it takes time to implement announced measures: in the case of the U.S., it took more than one year between the first plans for fiscal stimulus (cf. Summers, 2008) and the enactment of the American Recovery and Reinvestment Act of 2009. Our results suggest that an announced but delayed stimulus of this type can appear to be effective once it occurs, through lower credit spreads and higher output growth. However, the announcement of the stimulus can deepen the crisis before the implementation of any measures. Moreover, after its direct demand effects are realized, the stimulus tends to be followed by lower medium-term growth than without any government interventions.

Our findings thereby support the notion that it is important that fiscal stimulus is timely but with the warning that even a timely stimulus can have undesirable crowding-out effects that can potentially offset its desired impact. This warning, which hinges on the degree of intermediary financing, should be taken into account together with warnings on potential long-run crowding-out effects of fiscal stimulus e.g. due to distortionary taxation (see e.g. Coenen et al, 2010; Drautzburg and Uhlig, 2011).

If not fiscal stimulus, what can governments do to stabilize the economy in times of financial stress? Our results tend to confirm the conclusion from previous studies that financial sector policies can be effective tools. However, once realistic aspects of deficit financing are taken into account, such policies seem less effective. According to our findings, if transfers to intermediaries are not the favored means, an important condition for alternative means such as loans to intermediaries to be effective is that repayments are agreed to occur with a significant delay. If they need to be repaid relatively early, the funding pressures that are put on financial intermediaries by an increase in government borrowing can again offset the desired recapitalization effects. Remarkably, however, short-run stabilization gains are still possible even if loans are to be paid back at penalty rates. The benefits to intermediaries from such policies are then realized at the peak of the crisis and repayment effects tend to be discounted in a way that the initial stabilization gains are enhanced. Quite oddly, however, a similar model as ours that does not consider deficit financing of government policies predicts that very large penalty factors of more than two hundred percent can bring overall stabilization gains. The latter underlines the relevance of taking the consequences of
intermediary financing of government policies into account.

To conclude, we would like to emphasize that the model discussed in this chapter highlights specific mechanisms that seem relevant for an analysis of government policies in times of financial stress, while neglecting other well-known aspects that are relevant to characterize the macroeconomic effects of fiscal policy. We could look at additional policies such as tax cuts, transfers to liquidity-constrained consumers, public investment, labor market policies etc. in the context of our model. The effects of such policies are however well-studied by now, and extensions of our model into that direction are unlikely to change our main conclusions or to add much understanding to the key mechanisms highlighted in this chapter.

4.A Steady state solution

This appendix derives the solution for the non-stochastic steady state of the model and shows that the incentive constraint is binding in the steady state. For simplicity, the solution is derived for a zero inflation steady state. This is achieved by setting the target inflation rate in the monetary policy rule accordingly, i.e. $\bar{\pi} = 1$; the Taylor rule then implies that $\pi = 1$. The steady state real interest rate on deposits and the steady state risk-free nominal interest rate then follow from the household’s consumption Euler equation and the corresponding Fisher relation:

$$r^d = \beta^{-1} - 1, \quad r^n = r^d.$$ 

Further, by the capital producer’s first-order condition, the relative price of capital equals one in the steady state: $q = 1$.

To solve for the variables that are determined by the financial intermediaries’ problem, we guess and verify that there is an equilibrium with $r^k - r^d = r^b - r^d = \Gamma > 0$. We also take as given the total leverage ratio $\phi$ by calibrating $\chi$, the average survival time of bankers $\Theta = 1/(1 - \theta)$ by setting $\theta = (\Theta - 1)/\Theta$ and the interest rate spread $\Gamma$ by calibrating $\lambda$. As $r^k = r^b$, we then obtain from the portfolio manager’s first-order condition that $\omega = \bar{\omega}$. Given $r^d$, we also obtain $r^k = r^d + \Gamma$ and $r^b = r^k$. From the
equation for $r^p$, it follows that $r^p = r^k$. We further obtain

$$\varrho = 0, \quad v = \frac{\beta(1 - \theta)(1 + r^p)}{1 - \beta \theta}, \quad \eta = \frac{\beta(1 - \theta)(1 + r^d)}{1 - \beta \theta}, \quad \lambda = v + \frac{(1 - \phi)\eta}{\phi}.$$  

We also have

$$\chi = 1 - \theta(\Gamma \phi + 1 + r^d).$$

Next, we see that the incentive constraint indeed binds in the steady state, because

$$\lambda - v + \eta = \frac{\eta}{\phi} = (1 - \theta)\beta(1 + r^d)\phi^{-1}(1 - \theta \beta)^{-1} > 0.$$  

We now solve for the production allocation. From the price setting equations, for a zero inflation steady state, we have

$$\pi^\star = \Delta = 1, \quad \Xi_1 = m\lambda y(1 - \beta \psi)^{-1}, \quad \Xi_2 = \lambda y(1 - \beta \psi)^{-1},$$

such that $\Xi_1/\Xi_2 = m$. The first-order condition of the intermediate goods firms’ price-setting problem therefore implies that $m = (\epsilon - 1)/\epsilon$. As $\Delta = 1$ and $a = \xi = 1$, we will use that steady state final output is $y = k^\alpha h^{1-\alpha}$. Further, the steady state real wage can be derived from the marginal cost equation, given $r^k$ and $m$:

$$w = [\alpha^\alpha(1 - \alpha)^{1-\alpha}m(r^k + \delta)^{-\alpha}]^{1/\alpha}.$$  

The capital-labor ratio is then

$$k/h = \alpha(1 - \alpha)^{-1}w(r^k + \delta)^{-1}.$$  

By the resource constraint, the steady state ratio of consumption over output is

$$c/y = 1 - i/y - g/y,$$

where $i/y$ and $g/y$ are taken as given. The household’s remaining first-order conditions
for consumption and hours worked then imply that

\[
\lambda = (1 - \beta v)\{(1 - v)(c/y)y\}^{-1}, \quad h = \{(1 - \beta v)w\{(1 - v)(c/y)y\}^{-1}\}^{\frac{1}{\phi}}. \quad (4.27)
\]

Steady state final output then follows from \( y = (k/h)^{\alpha} h \), or

\[
y = (k/h)^{\frac{\alpha}{1+\alpha}}\{(1 - \beta v)w\{(1 - v)(c/y)y\}^{-1}\}^{\frac{1}{1+\alpha}}.
\]

such that \( \lambda \) and \( h \) can be computed from (4.27). Steady state consumption, investment, and government spending are thus

\[
c = (c/y)y, \quad i = (i/y)y, \quad g = (g/y)y.
\]

The government spending process can then be specified such that \( g/y \) can be taken as given, as it was assumed above, by setting \( \bar{g} = g \). The capital accumulation equation furthermore implies that \( i/k = \delta \). Steady state investment therefore satisfies \( i = \delta(k/h)h \). The steady state ratio of investment over GDP is thus

\[
i/y = \delta(k/h)(h/y) = \delta(k/h)^{1-\alpha} = \delta[\alpha(1 - \alpha)^{-1}w(rk + \delta)^{-1}]^{1-\alpha} = \delta\alpha m(rk + \delta)^{-1}.
\]

Solving the last equation for \( \delta \) yields

\[
\delta = r^k (i/y) (\alpha m - i/y)^{-1}.
\]

Hence, \( \delta \) can be calibrated such that \( i/y \) can be taken as given, as it was assumed above. The steady state capital stock then follows from the capital accumulation equation: \( k = i/\delta \). Given \( k \), we obtain the steady state level of claims on non-financial firms by financial intermediaries from the market clearing condition: \( s^k = k \). On the fiscal side, we take the steady state ratio of government debt over GDP \( b/y \) as given by calibrating the steady state level of taxes \( \bar{\tau} \), such that

\[
b = (b/y)y, \quad \bar{\tau} = g + r^b b.
\]
To equalize the demand for government bonds by financial intermediaries $s_b$ and bond supply by the government $s$, given $s_k$, we calibrate $\bar{\omega}$ accordingly, as $\omega = \bar{\omega}$ and $s_b/(1 - \omega) = \phi n = s_k/\omega$, or
\[
\omega = (s_k/s_b)(1 + s_k/s_b)^{-1}.
\]
Given $b$, we thus obtain the steady state level of the intermediaries’ government bond holdings from the market clearing condition: $s_b = b$. The remaining financial variables then follow as
\[
n = s_k(\omega \phi)^{-1}, \quad p = \phi n, \quad d = p - n.
\]

4.B The model without financial intermediaries

This appendix describes the version of the model without financial intermediaries. In this model, there are no bankers and households are thus formed entirely by infinitely lived workers with mass unity. Households save by investing in government bonds and by purchasing claims issued by intermediate goods firms. Accordingly, the budget constraint of a representative household becomes
\[
c_t + s_t^b + q_t s_t^k + \tau_t \leq w_t h_t + (1 + r_t^b) s_{t-1}^b + (1 - \bar{\tau}^k) (1 + r_t^k) q_{t-1} s_{t-1}^k + \Sigma_t.
\]
We have introduced a flat-rate tax on capital income $\bar{\tau}^k$ whose function is discussed below. With $\Lambda_{t,t+1}$ as in the main text, the first-order conditions for the household’s choices of $s_t^b$ and $s_t^k$ are
\[
s_t^b : 1 = \beta E_t \Lambda_{t,t+1} (1 + r_{t+1}^b), \quad (4.28)
\]
\[
s_t^k : 1 = \beta E_t \Lambda_{t,t+1} (1 - \bar{\tau}^k) (1 + r_{t+1}^k). \quad (4.29)
\]
As in the main text, the monetary authority sets the risk-free nominal interest rate $r_t^n$.

A Fisher relation defines the ex-post gross real interest rate on government bonds:
\[
1 + r_t^b = (1 + r_{t-1}^n) \pi_t^{-1}. \quad (4.30)
\]
On the fiscal side, the government budget constraint becomes

\[ b_t + \tau_t + \bar{\tau}^k(1 + r_t^k)q_{t-1}s_{t-1}^k = g_t + (1 + r_t^b)b_{t-1}. \]  

(4.31)

The rational expectations equilibrium of this model is then the set of sequences \( \{c_t, h_t, w_t, i_t, k_t, q_t, y_t, m_t, \pi_t, \pi^*_t, \Xi_1, \Xi_2, \Delta_t, r_t^k, r_t^b, s_t^k, s_t^b, b_t\}_{t=0}^{\infty} \) and shadow prices \( \{\lambda_t\}_{t=0}^{\infty} \), such that for given initial prices and initial values, a fiscal policy \( \{g_t, \tau_t\}_{t=0}^{\infty} \), a monetary policy \( \{r_t^m\}_{t=0}^{\infty} \), and sequences of shocks \( \{a_t, \xi_t\}_{t=0}^{\infty} \), conditions (4.1)-(4.2), (4.10)-(4.14), (4.19)-(4.20), (4.23)-(4.26), (4.28)-(4.31), and the transversality conditions are satisfied. The tax \( \bar{\tau}^k \) is calibrated such that this model implies the same steady state values for the relevant variables as the model with financial intermediaries, to make the impulse responses shown in the main text comparable. In particular, as \( r_t^b \) replaces the deposit rate \( r_t^d \) in this model, without any adjustment there would be no steady spread between \( r_t^k \) and \( r_t^b \). This means that steady state capital, investment, output, etc. would be higher than in the model with financial intermediation. To address this issue, we take a steady state that satisfies \( r^k = r^b + \Gamma \), with \( \Gamma > 0 \) as in the main text and \( r^b = r^m = \beta^{-1} - 1 \), and we calibrate \( \bar{\tau}^k \) to generate this spread. In particular, (4.28)-(4.29) imply that

\[ 1 + r^b = (1 - \bar{\tau}^k)(1 + r^k) = (1 - \bar{\tau}^k)(1 + r^b + \Gamma), \]

or

\[ \bar{\tau}^k = \Gamma(1 + r^b + \Gamma)^{-1} > 0. \]

For a small spread, \( \bar{\tau}^k \) will be small enough not to have a significant impact on the dynamics. In addition, we need to change the calibration of \( \tau \) to have identical values for the fiscal variables:

\[ \bar{\tau} = g + r^b b - \bar{\tau}^k(1 + r^k)s^k. \]

The steady state calculations for the remaining relevant variables and parameters are identical to those described in Appendix 4.A.
Chapter 5

Sovereign Default Risk and Macroeconomic Fluctuations in an Emerging Market Economy*

Abstract

This chapter examines the role of sovereign default risk for business cycle dynamics in an emerging market economy. We assess whether a small open economy model with sticky prices and endogenous sovereign default premia can explain the fluctuations of inflation, interest rates, and government debt along with real business cycles in Turkey. We compare estimation results for a basic version of the model where government debt is only relevant for the dynamics of fiscal variables and an augmented version where government debt matters for the dynamics of all variables due to a perceived risk of sovereign debt default. The results show that, unlike the basic model, the augmented model does not rely on large shocks or extreme parameter values to match the data. Comparatively small shocks are instead amplified through a self-enforcing feedback cycle between government debt, default premia, and nominal variables which brings the model closer to the data.

5.1 Introduction

A growing empirical literature seeks to explain business cycles in developed countries using new open-economy macroeconomic (NOEM) models (e.g. Adolfson, Laséen, Lindé, and Villani, 2007; Justiniano and Preston, 2008, 2010; Lubik and Schorfheide, 2007). The evidence for emerging market countries is however still scarce. One possible

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*This chapter is based on joint work with Malte Rieth.
reason is that emerging market economies are often characterized by business cycle fluctuations that are difficult to generate with standard NOEM models. In particular, many emerging countries such as Argentina, Brazil, Mexico, Russia, and Turkey have been characterized by highly persistent and volatile inflation, nominal interest rates, and government debt. This chapter argues that it is useful to model investors’ beliefs on sovereign debt default in order to obtain such dynamics in an empirical NOEM model for an emerging market economy.

We set up a small open economy model with sticky prices following Galí and Monacelli (2005) but including a government which borrows in domestic currency at home and in foreign currency abroad in line with the “original sin” literature (Eichengreen and Hausmann, 1999). The government follows a tax rule with at least some feedback from higher debt levels on taxes. Following Schabert and van Wijnbergen (2011), we argue that this rule may imply perceived infeasible rates of taxation, in which case the government is expected to default on (part of) its outstanding debt. The presence of sovereign default beliefs introduces an endogenous default premium, which depends on real government liabilities, in the households’ consumption Euler equation. If the monetary authority follows an active interest rate policy, inflation increases imply higher real rates and an associated increase in the public debt service burden, which may lead to sovereign default fears. The latter generates a negative feedback from government debt on its expected return, which implies that current savings tend to be lower, putting pressure on the exchange rate, which increases inflation and the need for the monetary authority to raise interest rates, further increasing debt obligations and default premia, and so on. This destabilizing effect of active monetary policy in the presence of sovereign default beliefs, which has been pointed out by Blanchard (2005) and which has been analyzed theoretically by Schabert and van Wijnbergen (2006, 2011), is at the heart of our model.

Using this model we assess the role of sovereign default risk in explaining business cycle fluctuations in an emerging market economy, taking Turkey’s experience as a natural experiment. In particular, Turkey was hit by a severe financial crisis in November 2000 when nominal interest rates increased sharply, accompanied by a downgrading of government debt to below investment grade. These observations indicate that fears of
sovereign default played a relevant role, although a debt default did not actually occur. We thus estimate two variants of the model on quarterly Turkish data for the period 1994Q3-2008Q2 by Bayesian methods. The two variants differ only with respect to the existence of the default premium in the consumption Euler equation. Without that premium the model reduces to a standard NOEM model where the level of government debt is only relevant for the level of (lump-sum) taxes. With the premium the level of government debt matters for the dynamics of all variables.

Our results show that the basic model without sovereign default risk relies on relatively large shocks and extreme parameter values to match the data. The augmented model with default risk neither requires very large shocks nor extreme parameter values; instead, comparatively small shocks are amplified and propagated through the feedback mechanism between government debt, default premia, and nominal variables described above. Formal model comparisons clearly support the modification of the consumption Euler equation in the augmented model, whose forecasting performance is also significantly improved. We therefore conclude that modelling investors’ beliefs on sovereign debt default can indeed lead to a better understanding of business cycle fluctuations in an emerging market economy.

In addition to the above-cited studies on empirical NOEM models, our study is related to the following literature. The empirical shortcomings of the consumption Euler equation have led some researchers to include reduced-form risk premium shocks in both closed and open economy models (e.g. Adolfson et al., 2007; Justiniano and Preston, 2010; Smets and Wouters, 2007). Other studies have explored the role of permanent productivity shocks (Aguiar and Gopinath, 2007) and financial frictions (e.g. Chang and Fernández, 2010; García-Cicco, Pancraci, and Uribe, 2010; Neumeyer and Perri, 2005; Uribe and Yue, 2006) in explaining the empirical regularities of emerging market business cycles, in particular the high volatility of consumption relative to output and the countercyclicality of interest rates. Based on the seminal contribution of Eaton and Gersovitz (1981), the role of strategic government default has been investigated by Arellano (2008). She focuses on the terms of international loans that are endogenous to domestic fundamentals and depend on the incentives to default in order to explain co-movements between real interest rates and output.
Inspired by the idea that financial frictions are important for understanding emerging market business cycles, our contribution is to focus on the role of fiscal sustainability concerns in a model with nominal rigidities. While we also use stochastic shocks to describe past business cycles in Turkey, our findings suggest that a standard set of shocks cannot explain the Turkish experience in an NOEM model where fiscal sustainability concerns are absent. Instead, we find that the amplification and propagation mechanism due to sovereign risk does seem to be an important driving force of business cycle fluctuations in Turkey.\footnote{Given the findings of Chang and Fernández (2010) and García-Cicco, Pancrazi, and Uribe (2010), indicating that permanent productivity shocks tend to have a negligible role in explaining business cycles in emerging economies compared to financial frictions, we neglect such shocks in our analysis.} Finally, as in Schabert and van Wijnbergen (2011), in our model there is no strategic motive for the government to default on its debt. Default premia are instead determined by investors’ beliefs that infeasible rates of taxation may force the government into default.

The remainder of the paper is organized as follows. Section 5.2 lays out the model. Section 5.3 describes the estimation of the model on quarterly Turkish data from 1994Q3 to 2008Q2. Section 5.4 presents the estimation results. We compare the basic model and the augmented model in terms of parameter estimates, marginal data densities, variance decompositions, and forecasting performance. We then implement counterfactual experiments based on the augmented model to understand how business cycles in Turkey have been influenced by sovereign risk, before comparing the impulse response dynamics generated by both model versions. Finally, we briefly summarize various sensitivity checks. Section 5.5 concludes.

5.2 Model description

We outline a small open economy model with sticky prices based on Galí and Monacelli (2005). The model considers expectations of sovereign default following Schabert and van Wijnbergen (2011), whose presence breaks Ricardian equivalence since the time path of government debt matters for equilibrium determination, but the model nests the Ricardian case where the time path of government debt is irrelevant for the allocation. We allow for foreign currency denominated debt to provide a realistic description of
the conduct of fiscal policy in Turkey, where the government can only borrow limited amounts from abroad in Turkish lira. Eichengreen and Hausmann (1999) call this the “original sin” which typically characterizes emerging market economies. Changes in the real exchange rate then have a direct impact on expected sovereign default rates due to the presence of foreign currency denominated debt.

5.2.1 Public sector

The public sector consists of a government and a monetary authority. The price of domestic bonds is set by the monetary authority, and since government bonds are subject to perceived default risk, the monetary policy instrument is an interest rate on an asset which exhibits a contingent pay-off. Thus, the policy instrument carries a risk component that will be reflected in equilibrium (see Blanchard, 2005; Loyo, 2005; Schabert and van Wijnbergen, 2006, 2011).

Fiscal policy

The government issues one-period discount bonds denominated in domestic and foreign currency $B_{H,t}$ and $B_{F,t}$, respectively. It levies lump-sum taxes $P_t \tilde{\tau}_t$ on domestic households and it purchases domestic goods $P_{H,t}g_t$, where $P_t$ and $P_{H,t}$ denote the consumer price level and the price of domestically produced goods, respectively. The monetary authority sets the domestic currency price $1/R_{H,t}$ of domestic bonds, whereas the foreign currency price $1/R_{F,t}$ of foreign currency denominated bonds is endogenously determined in equilibrium.

The government is assumed to follow a simple tax feedback rule, adjusting lump-sum taxes in response to the outstanding stock of debt:

$$P_t \tilde{\tau}_t = \kappa (B_{H,t-1} + X_t B_{F,t-1}) + P_t \exp(\varepsilon_{\tau,t}), \quad (5.1)$$

$^2$Time is indexed by $t = 0, 1, 2, \ldots, \infty$. Throughout, nominal (real) variables are denoted by capital (lower) letters, asterisks denote foreign variables and variables without time subscript denote non-stochastic steady state values.

$^3$The assumption that government purchases are fully allocated to domestically produced goods is motivated by empirical evidence for OECD countries of a strong home bias in government procurement, above that observed for private consumption (see e.g. Trionfetti, 2000; Brulhart and Trionfetti, 2004).
where $\varepsilon_{t,t} \sim N(0, \sigma^2_\varepsilon)$ is a lump-sum tax shock and $X_t$ denotes the domestic currency price of one unit of foreign currency. Following Bohn (1998), a tax rule of this type ensures fiscal solvency for any finite initial level of debt as long as $\kappa > 0$. However, it may imply politically infeasible levels of taxation as we discuss next.

Following Schabert and van Wijnbergen (2011), according to investors’ beliefs, the government defaults when debt service would demand a politically infeasible level of taxation $T$. Lenders do not know the exact value of $T$, but they have a prior on its distribution, $h(T)$. Given that tax revenues are set according to (5.1), the perceived probability of default $\delta_t$ then equals the probability that the tax rule implies a level of $\tilde{\tau}_t$ exceeding $T$:

$$\delta_t = \int_0^{\tilde{\tau}_t} h(T) dT.$$ (5.2)

Let $q_t = X_t P_t^*/P_t$ denote the real exchange rate and $\pi_t = P_t/P_{t-1}$ and $\pi_t^* = P_t^*/P_{t-1}$ home and foreign consumer price index (CPI) inflation, respectively. For a differentiable distribution function $h(\cdot)$ the impact of total real debt $b_t = b_{H,t-1} \pi_t^{-1} + q_t b_{F,t-1} \pi_t^* \pi^{-1}_t$ on the probability of default is given by $\partial \delta_t(\cdot)/\partial b_t = \kappa h(\kappa b_t) > 0$. Thus, the perceived default probability strictly increases with the real value of total debt. For the local analysis of the model we use the product of the ratio $(b_H/\pi)/(1 - \delta)$ and the elasticity of the default probability with respect to the real value of total debt evaluated at the steady state:

$$\Phi = \frac{b_H/\pi}{1 - \delta} \left. \frac{\partial \delta_t(\cdot)}{\partial b_t} \right|_{b_t = b},$$

where $\delta = \delta(b) < 1$. We refer to $\Phi$ as the default elasticity and we treat it as a structural parameter in the empirical implementation. Note that $\Phi > 0$ if $b_H/\pi > 0$.

To determine the division of total debt among domestic debt and foreign debt, we assume that the government issues foreign currency denominated debt as a time-varying fraction $f_t \geq 0$ of domestic debt, $X_t B_{F,t}/R_{F,t} = f_t B_{H,t}/R_{H,t}$, where $f_t$ follows an autoregressive process in logs: $\log(f_t/\bar{f}) = \rho_f \log(f_{t-1}/\bar{f}) + \varepsilon_{f,t}$, with $\rho_f \in [0, 1)$ and $\varepsilon_{f,t} \sim N(0, \sigma^2_{\varepsilon_f})$. It is also assumed that the savings through default, $\delta_t (B_{H,t-1} + X_t B_{F,t-1})$, are handed out in a lump-sum fashion to domestic households. Given the

\footnote{As noted by Schabert and van Wijnbergen (2011), the structure of the default premium has a structure that commands broad empirical support: see, for instance, Cantor and Packer (1996), Edwards (1994), Eichengreen and Mody (2010), Ferucci (2003), and Min (1998)}.
specification (5.2), the period-by-period perceived government budget constraint for any period $t$ reads as follows:

$$B_{H,t}/R_{H,t} + X_t B_{F,t}/R_{F,t} + P_t \tau_t = P_{H,t} g_t + (1 - \delta_t)(B_{H,t-1} + X_t B_{F,t-1}),$$

where $P_t \tau_t = P_t \tilde{\tau}_t - \delta_t(B_{H,t-1} + X_t B_{F,t-1})$ and $g_t$ follows an autoregressive process in logs: $\log(g_t/\bar{g}) = \rho_g \log(g_{t-1}/\bar{g}) + \varepsilon_{g,t}$, with $\rho_g \in [0, 1)$ and $\varepsilon_{g,t} \sim N(0, \sigma_g^2)$.

### Monetary policy

In line with the actual behavior of the Central Bank of the Republic of Turkey (CBRT), CPI inflation stabilization is assumed to be the target of monetary policy. The monetary authority thus sets the domestic currency price of domestic bonds according to the reaction function

$$R_{H,t}/R_H = (\pi_t/\pi)^{\alpha_\pi} \exp(\varepsilon_{R,t}),$$

where $\varepsilon_{R,t} \sim N(0, \sigma_R^2)$. Interest rate smoothing did not seem to be a primary goal of the CBRT, so we do not include a smoothing term in the reaction function. As we are primarily interested in the interaction between an inflation targeting monetary authority and fiscal policy, we also do not include an output term for now. The CBRT was however targeting the exchange rate before the economic reforms in 2001 (see Gormez and Imaiz, 2007). We therefore check the sensitivity of the estimation results to adding an exchange rate term and also an output term to the reaction function in Section 5.4.4.

### 5.2.2 Private sector

### Domestic households

The domestic economy is inhabited by a continuum of infinitely lived households with identical asset endowments and identical preferences. A representative domestic household chooses consumption $c_t$, hours worked $n_t$, and the asset portfolio described below,
to maximize

\[ E_t \sum_{s=0}^{\infty} \beta^s \left[ z_{t+s} \frac{1}{1-\sigma} c_{t+s}^{1-\sigma} - \frac{1}{1+\eta} n_{t+s}^{1+\eta} \right], \quad \beta \in (0,1), \quad \sigma > 0, \quad \eta \geq 0, \quad (5.4) \]

where \( z_t \) is a demand shock which follows an autoregressive process in logs: \( \log z_t = \rho z_{t-1} + \varepsilon_{zt}, \) with \( \rho \in (0,1) \) and \( \varepsilon_{zt} \sim N(0,\sigma^2_z). \) Domestic households invest in domestic and foreign currency denominated government bonds and in a complete set of state-contingent securities which are traded internationally. Let \( \Gamma_{t,t+1} \) denote the stochastic discount factor for a one-period ahead nominal payoff \( S_{t+1} \) in foreign currency. The perceived flow budget constraint, which takes into account the household’s default beliefs, is given by

\[
P_t c_t + P_t \tau_t + E_t(X_t \Gamma_{t,t+1} S_{t+1}) + B_{H,t} / R_{H,t} + X_t B_{F,t}^b / R_{F,t} \leq X_t S_t + (1 - \delta_t)(B_{H,t-1} + X_t B_{F,t-1}^b) + P_t w_t n_t + \Sigma_t, \quad (5.5)
\]

for given initial wealth endowments \( B_{H,-1}, B_{F,-1}^b, \) and \( S_0. \) Here, \( w_t \) is the real wage rate and \( \Sigma_t \) collects payouts from ownership of firms, which are both taken as given by the household.

The household’s consumption basket is an aggregate of domestically produced goods \( c_{H,t} \) and goods of foreign origin \( c_{F,t}, \) \( c_t = \gamma c_{H,t}^{1-\vartheta} c_{F,t}^\vartheta, \) where \( \vartheta \in [0,1] \) denotes the import share and \( \gamma = \vartheta^{-\vartheta} (1 - \vartheta) \vartheta^{-1}. \) The optimal allocation of consumption among \( c_{H,t} \) and \( c_{F,t} \) yields the demand functions \( c_{H,t} = (1 - \vartheta) P_t c_t / P_{H,t} \) and \( c_{F,t} = \vartheta P_t c_t / P_{F,t}, \) where \( P_{F,t} \) is the price of foreign goods. The definition of the CPI follows as \( P_t = P_{H,t}^{1-\vartheta} P_{F,t}^\vartheta. \) The first-order conditions from maximization of (5.4) subject to a no-Ponzi-game condition and (5.5) are as follows:

\[
\lambda_t = z_t c_t^{-\sigma}, \quad \lambda_t = R_{H,t} \beta E_t [(1 - \delta_{t+1}) \lambda_{t+1} \pi_{t+1}^{-1}], \\
n_{t}^{\eta} = \lambda_t w_t, \quad \lambda_t q_t = R_{F,t} \beta E_t [(1 - \delta_{t+1}) \lambda_{t+1} q_{t+1} \pi_{t+1}^{-1}], \\
\Gamma_{t,t+1} = \beta (X_{t+1} / X_t)(\lambda_{t+1} / \lambda_t) \pi_{t+1}^{-1}, \quad (5.6)
\]

where \( \lambda_t \) denotes the Lagrangian multiplier associated with (5.5). The budget con-
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straint holds with equality and the transversality conditions are satisfied. It follows that, everything else equal, a higher expected default rate leads households to demand a higher interest rate.

**Foreign households**

The foreign economy is inhabited by a continuum of infinitely lived households with identical asset endowments and which have the same preference structure as domestic households. A representative foreign household’s demand for domestically produced goods $c_{H,t}^*$ satisfies

$$c_{H,t}^* = \vartheta^* P_t^* c_t^*/P_{H,t}^*,$$

(5.7)

where $\vartheta^* \in [0, 1]$, $c_t^*$ is aggregate foreign consumption, and $P_t^*$ is the price of domestic goods expressed in foreign currency. The foreign household invests in state-contingent securities $S_t$ and foreign currency denominated bonds issued by the domestic government $B_{F,t}^F$. The first order conditions are given by

$$\lambda_t^* = R_{F,t} \beta E_t[(1 - \delta_{t+1}) \lambda_{t+1}^* \pi_{t+1}^* - 1]$$

and

$$\Gamma_{t,t+1} = \beta (\lambda_{t+1}^* / \lambda_t^*) \pi_{t+1}^* - 1,$$

(5.8)

where $\lambda_t^* = c_t^*/\pi_t^*$. As the foreign economy is exogenous to the domestic economy, we assume for simplicity that foreign consumption and inflation follow a VAR process in logs with $p$ lags: 

$$\log(c_t^*/\bar{c}^*), \log(\pi_t^*/\bar{\pi}^*)' = \sum_{i=1}^p \Phi_{ts}[\log(c_{t-i}^*/\bar{c}^*), \log(\pi_{t-i}^*/\bar{\pi}^*)]' + [v_{c,t}, v_{\pi,t}]'$$

where $I - \sum_{i=1}^p \Phi_{ts}$ is a non-singular matrix and $[v_{c,t}, v_{\pi,t}]' \sim N(0, \Sigma_s)$. Notice also that $B_{F,t}^F + B_{F,t}^T = B_{F,t}$, for all $t$, by market clearing.

**Production and pricing**

The production sector consists of final goods producers and intermediate goods producers. Final goods producers are perfectly competitive. They assemble the final domestic good $y_{H,t}$ from intermediate goods $y_{i,t}^H$, $i \in [0, 1]$, through the technology

$$y_{H,t} = \int_0^1 (y_{H,t}^i)^{(\epsilon-1)/\epsilon} di / (\epsilon-1),$$

where $\epsilon$ denotes the elasticity of substitution among intermediate goods. Taking as given all intermediate goods prices $P_{H,t}^i$ and the final
goods price $P_{H,t}$, profit maximization yields input demands

$$y_{H,t}^i = (P_{H,t}^i/P_{H,t})^{-\varepsilon} y_{H,t} \tag{5.9}$$

for all $i$, where we have used the zero profit condition in the final goods sector, i.e. $P_{H,t}y_{H,t} = \int_0^1 P_{H,t}^i y_{H,t}^i di$. The price index for domestic goods $P_{H,t}$ follows from using (5.9) in the zero profit condition stated above: $P_{H,t} = \left[\int_0^1 (P_{H,t}^i)^{1-\varepsilon} di\right]^{1/(1-\varepsilon)}.$

Intermediate goods production is conducted by a continuum of monopolistically competitive firms. Each firm $i$ uses the technology $y_{H,t}^i = a_t n_i$, where $a_t$ is common total factor productivity which follows an autoregressive process in logs: $\log a_t = \rho_a \log a_{t-1} + \varepsilon_{a,t}$, with $\rho_a \in [0, 1)$ and $\varepsilon_{a,t} \sim N(0, \sigma_a^2)$. Intermediate goods producers solve a two-stage problem. In the first stage, taking the input price $w_t$ and the output price $P_{i,H,t}$ as given, firms hire labor to minimize costs, which yields the first-order conditions $P_{i,t} w_t = MC_{it} a_t$ for all $i$, where $MC_{it}$ denotes nominal marginal costs. The equality $MC_{it} = MC_t$ holds since all firms face the same input prices and use the same technology. Expressing real marginal costs in terms of domestic prices, $mc_t = MC_t/P_{H,t}$, then yields the labor demand function $w_t = mc_t a_t P_{H,t}/P_t$.

In the second stage, given real marginal costs, intermediate goods producers choose prices $P_{i,H,t}$ to maximize profits. We allow for staggered price setting following Calvo (1983) and Yun (1996). Each period a fraction $1 - \phi$ of randomly selected firms is allowed to set an optimal new price $\hat{P}_{H,t}$ or $\hat{P}_{H,t}$, by symmetry. The remaining firms adjust their prices along with steady state producer price inflation $\pi_H$. Each firm $i$ which receives permission to optimally reset its price maximizes the expected sum of discounted profits subject to (5.9):

$$\max_{E_t} \sum_{s=0}^{\infty} \phi^s X_t \Gamma_{t,t+s}(P_{H,t}^{i,s} - P_{H,t+s}mc_{t+s}) y_{H,t+s}^i \quad \text{s.t.} \quad y_{H,t+s}^i = (P_{H,t+s}^i/P_{H,t+s})^{-\varepsilon} y_{H,t+s},$$

where $P_{H,t+s} = \hat{P}_{H,t} \pi_{H,t}^s$ for $s = 1, 2, \ldots, \infty$. The first-order condition is given by

$$0 = E_t \sum_{s=0}^{\infty} \phi^s X_t \Gamma_{t,t+s} y_{H,t+s}^i [(1 - \varepsilon) \pi_{H,t}^s \hat{P}_{H,t} + \varepsilon P_{H,t+s}mc_{t+s}].$$
The price index of domestic goods follows as $P_{H,t}^{1-\epsilon} = (1 - \phi)\tilde{P}_{H,t}^{1-\epsilon} + \phi(P_{H,t-1})^{1-\epsilon}$.

### 5.2.3 Market clearing

Market clearing requires that the demand for labor services is equal to labor supply: $\int_0^1 n_i di = n_t$. Integrating $y_{H,t} = a_i n_i$ over all $i$, it then follows that $\int_0^1 y_{H,t} di = a_t n_t$. We assume that the domestic economy is small relative to the foreign economy, which implies that the foreign producer price level $P_{F,t}^*$ is identical to the foreign consumption price index $P_{F,t}^*$. Furthermore, the law of one price is assumed to hold separately for each good, such that $P_{F,t} = X_t P_{F,t}^*$ and $P_{H,t} = X_t P_{H,t}^*$. Using the definition of the CPI, foreign demand for domestic goods (5.7) can then be re-written as $c_{H,t}^* = \vartheta q_{t}^{1/(1-\vartheta)} c_t^*$ and domestic demand $c_{H,t} = (1 - \vartheta)P_t c_t/P_{H,t}$ can be re-written as $c_{H,t} = (1 - \vartheta)q_t^{\vartheta/(1-\vartheta)} c_t$, where we have used that $P_{H,t}/P_t = q_t^{\vartheta/(\vartheta-1)}$. Goods market clearing requires that aggregate supply $y_{H,t}$ equals aggregate demand $c_{H,t} + c_{H,t}^* + g_t$. The goods market clearing condition can be re-written as follows:

$$y_{H,t} = (1 - \vartheta)q_t^{\vartheta/(1-\vartheta)} c_t + \vartheta q_t^{1/(1-\vartheta)} c_t^* + g_t.$$

The CPI inflation rate can be expressed in terms of producer price inflation through $\pi_t = \pi_{H,t}(q_t/q_{t-1})^{\vartheta/(1-\vartheta)}$ for all $t \geq 1$. Finally, combining (5.6) and (5.8) yields the international risk sharing condition $\lambda_t^* = \xi q_t \lambda_t$, which determines the relation between the levels of domestic and foreign marginal utility and the real exchange rate up to a positive constant $\xi$ that depends on initial endowments.

### 5.3 Model estimation

We employ a log-linear approximation of the system of equilibrium conditions around the non-stochastic steady state. Appendix 5.A provides the log-linearized system and a definition of the rational expectations equilibrium. The model is then estimated by Bayesian methods as described in An and Schorfheide (2007).
5.3.1 Methodology

Formally, let \( P(\theta_{M_i}|M_i) \) denote the prior distribution of the vector of structural parameters \( \theta_{M_i} \) for model \( M_i \), and let \( L(Y^T|\theta_{M_i}, M_i) \) denote the likelihood function for the observed data \( Y^T = [Y_1, \ldots, Y_T]' \). For \( t = 1, \ldots, T \), the solution to the log-linearized model has a state-space representation with the state equation \( x_t = Fx_{t-1} + G\varepsilon_t \) and the observation equation \( Y_t = Hx_t + u_t \), where the vectors \( x_t, \varepsilon_t \sim N(0, \Sigma_\varepsilon) \) and \( u_t \sim N(0, \Sigma_u) \) collect model variables, structural shocks, and measurement errors, respectively. The Kalman filter is applied to evaluate \( L(Y^T|\theta_{M_i}, M_i) \) and the posterior distribution

\[
P(\theta_{M_i}|Y^T, M_i) = \frac{L(Y^T|\theta_{M_i}, M_i) P(\theta_{M_i}|M_i)}{\int L(Y^T|\theta_{M_i}, M_i) P(\theta_{M_i}|M_i) d\theta_{M_i}} \propto L(Y^T|\theta_{M_i}, M_i) P(\theta_{M_i}|M_i)
\]

is evaluated by the Random Walk Metropolis (RWM) algorithm. The evidence of model \( M_i \) over another (not necessarily nested) model \( M_j \) is assessed by the Bayes factor \( p(Y^T|M_i)/p(Y^T|M_j) \), which summarizes the sample evidence in favor of model \( M_i \), the marginal data density \( p(Y^T|M_i) = \int L(Y^T|\theta_{M_i}, M_i) P(\theta_{M_i}|M_i) d\theta_{M_i} \) indicating the likelihood of model \( M_i \) conditional on the observed data. Further, for \( t = 1, \ldots, T \), the shocks \( \varepsilon_{t|T} \) are recovered by an application of the Kalman smoother at the parameter estimates. This step also yields smoothed estimates \( x_{t|T} \) of the unobserved states. To evaluate the forecasting performance of alternative models, one-step ahead forecasts are computed conditional on period \( t \) information: \( Y_{t+1|t} = Hx_{t+1|t} \), for all \( t \), where \( x_{t+1|t} = Fx_{t|t} \) and \( x_{t|t} \) denote updates from the Kalman filter.\(^5\)

5We use version 4.2.0 of the Dynare toolbox for MATLAB in the computations. The marginal data densities are estimated using Geweke’s (1999) modified harmonic mean estimator.

5.3.2 Data description

We use quarterly Turkish data on real GDP, real private consumption, the annualized consumer price inflation rate, the nominal interest rate on 3-month Turkish lira denominated treasury bills, the real effective exchange rate, real government consumption, real Turkish lira denominated domestic government debt, real foreign consumption,
and the foreign consumer price inflation rate. The sample period is 1994Q3-2008Q2.

Foreign variables are computed as trade-weighted averages of data for the U.S. and the euro area, which are Turkey’s main trading partners. Nominal variables are demeaned in consistence with their steady state values. Real variables are transformed into natural logarithms and they are detrended using a linear trend. Further details on data sources, definitions, and the construction of the foreign variables are provided in Appendix 5.B.

Domestic and foreign inflation (INFₜ and INFₜ*) and the domestic interest rate (INTₜ) are related to the model variables through the measurement equations INFₜ = 4\bar{\pi} \hat{\pi}ₜ, INFₜ* = 4\bar{\pi}* \hat{\pi}ₜ*, and INTₜ = 4\bar{R}_H \hat{R}_H,t. As the real effective exchange rate (REERₜ) is constructed from all trading partners, it is not exactly equivalent to the model-implied real exchange rate, which relates to the U.S. and the euro area. To account for this fact, we include a measurement error in the equation for the real exchange rate REERₜ = \hat{q}ₜ + uₜ with uₜ \sim N(0,\sigma^2_u).

5.3.3 Calibration

The steady state values that matter for the dynamics are calibrated to match sample averages. To match the average annual Turkish CPI inflation rate over the sample period of 37.2 percent, we set \bar{\pi} = 1.093. The average annualized 3-month treasury bill rate was approximately 72.4 percent, so we set \bar{R}_H = \hat{R}_H = 1.181. The shares of private and government consumption in GDP and the share of foreign currency debt over domestic currency debt are also set to their empirical counterparts, i.e. s_c = \bar{s}_c = 0.683, s_g = \bar{s}_g = 0.108, and f = \bar{f} = 0.829. Furthermore, the subjective discount factor β is set to 0.99, which implies a steady state default probability δ = \bar{\delta} = 1 - \bar{\pi}/\bar{R}_H/β = 0.065. The latter agrees with the average J.P. Morgan Emerging Market Bond Index Global (EMBIG) spread on Turkish governments bonds.

The parameters of the stochastic process for the foreign variables are calibrated by

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6We only observe domestic currency denominated debt since the Turkish government issues external debt only at maturities longer than 3 months. Note that the nominal interest rate also refers to domestic debt.

7We have verified that our main results are robust when estimating the model on data which was detrended using linear-quadratic and Hodrick-Prescott filtered trends.
fitting an identified VAR with \( p = 4 \) lags to detrended foreign consumption and the demeaned annual foreign inflation rate.\(^8\) Steady state foreign inflation \( \pi^* = \bar{\pi}^* \) is calibrated to match an average annual inflation rate of 2.4 percent. Foreign consumption and inflation are then included in the actual estimation step (fixing the VAR parameters) in order to recover the shocks of foreign origin. Finally, as the inverse Frisch elasticity of labor supply \( \eta \) is difficult to identify, we set it to 2, in line with available estimates (see Christoffel, Coenen, and Warne, 2008).

5.3.4 Priors

As, to our knowledge, the present study is the first study that estimates a DSGE model for Turkey, we have little prior information on the model’s deep structural parameters. We therefore use uniform priors on theoretically plausible ranges for most parameters as we would with restricted maximum likelihood estimation. The prior distributions are provided in Table 5.1.

The inverse elasticity of intertemporal substitution \( \sigma \) and the inflation feedback in the monetary reaction function \( \alpha_\pi \) obtain a lower bound of 0 and upper bounds of 20 and 10, respectively. The Calvo probability \( \phi \) and the domestic degree of openness \( \vartheta \) are restricted to the range \([0, 1]\) in consistence with their feasible values. To ensure a positive default elasticity, which is the case if steady state domestic debt is positive, the debt response \( \kappa \) in the tax rule is restricted to be larger than \( \kappa_L = 1 - \beta(1 - \delta) \) with an upper bound of 10.\(^9\) We use uniform priors on the range \([0, 1]\) for the \( AR(1) \) coefficients of the stochastic processes. The standard deviations of the shocks turned out to be weakly identified especially for the model without default risk (for which \( \Psi \) is set to zero). We therefore elicit inverse gamma priors with mean 0.05 and an infinite standard deviation, implying that a larger portion of the probability mass tends to fall on existing estimates for small open economies (see e.g. Adolffson et al., 2007; Justiniano and Preston, 2010; Lubik and Schorfheide, 2007), while the distribution still covers all of the feasible range.

\(^8\)Our identifying assumption is that foreign consumption affects foreign inflation within a quarter but not vice versa, which is achieved through a recursive Cholesky identification scheme.

\(^9\)The steady state satisfies \( b_H / \pi = [g(1 + \bar{f})^{-1}]/[\kappa + \beta(1 - \delta) - 1] \) and therefore \( \Phi > 0 \) when \( b_H / \pi > 0 \), which is the case if \( \kappa > 1 - \beta(1 - \delta) \) since \( g, \bar{f} > 0 \).
Table 5.1: Prior distributions and posterior estimates for Turkey

<table>
<thead>
<tr>
<th>Parameter</th>
<th>Dom. Prior</th>
<th>With sov. risk (M₁)</th>
<th>No sov. risk (M₂)</th>
</tr>
</thead>
<tbody>
<tr>
<td>σ</td>
<td>Inv. subst. elast.</td>
<td>U(0.20)</td>
<td>1.87 [0.98, 2.81]</td>
</tr>
<tr>
<td>φ</td>
<td>Price stickiness</td>
<td>U(0.1)</td>
<td>0.46 [0.38, 0.55]</td>
</tr>
<tr>
<td>θ</td>
<td>Openness</td>
<td>U(0.1)</td>
<td>0.13 [0.05, 0.20]</td>
</tr>
<tr>
<td>Φ</td>
<td>Default elasticity</td>
<td>R⁺</td>
<td>0.29 [0.24, 0.33]</td>
</tr>
<tr>
<td>αₙ</td>
<td>Mon. infl. resp.</td>
<td>R⁺</td>
<td>1.83 [1.61, 2.05]</td>
</tr>
<tr>
<td>κ</td>
<td>Tax debt resp.</td>
<td>U(κ₀, 10)</td>
<td>0.59 [0.51, 0.66]</td>
</tr>
<tr>
<td>ρₜ</td>
<td>AR(1) demand</td>
<td>[0.1]</td>
<td>0.77 [0.62, 0.93]</td>
</tr>
<tr>
<td>ρₐ</td>
<td>AR(1) productivity</td>
<td>[0.1]</td>
<td>0.96 [0.92, 1.00]</td>
</tr>
<tr>
<td>ρ₉</td>
<td>AR(1) gov. cons.</td>
<td>[0.1]</td>
<td>0.44 [0.31, 0.58]</td>
</tr>
<tr>
<td>ρ₇</td>
<td>AR(1) debt share</td>
<td>[0.1]</td>
<td>0.90 [0.82, 0.98]</td>
</tr>
<tr>
<td>σ₉</td>
<td>Std. demand inn.</td>
<td>R⁺</td>
<td>0.05 [0.03, 0.08]</td>
</tr>
<tr>
<td>σₐ</td>
<td>Std. prod. inn.</td>
<td>R⁺</td>
<td>0.03 [0.02, 0.04]</td>
</tr>
<tr>
<td>σ₉</td>
<td>Std. gov. cons. inn.</td>
<td>R⁺</td>
<td>0.04 [0.03, 0.04]</td>
</tr>
<tr>
<td>σ₇</td>
<td>Std. debt share inn.</td>
<td>R⁺</td>
<td>0.25 [0.21, 0.29]</td>
</tr>
<tr>
<td>σ₉</td>
<td>Std. tax inn.</td>
<td>R⁺</td>
<td>0.07 [0.06, 0.09]</td>
</tr>
<tr>
<td>σ₉</td>
<td>Std. int. rate inn.</td>
<td>R⁺</td>
<td>0.06 [0.05, 0.07]</td>
</tr>
<tr>
<td>σ₉</td>
<td>Std. meas. REERₜ</td>
<td>R⁺</td>
<td>0.11 [0.09, 0.13]</td>
</tr>
<tr>
<td>σ₉</td>
<td>Corr. tax, gov. cons.</td>
<td>[-1,1]</td>
<td>0.79 [0.66, 0.92]</td>
</tr>
</tbody>
</table>

Log data density³ | 895.51 | 748.23

---

a The results are based on 500,000 draws from the RWM sampler, dropping the first 250,000 draws, and an average acceptance rate of approximately 25 percent. Posterior means are reported.

b U(a, b) refers to the continuous uniform distribution with lower bound a and upper bound b; IG(c) refers to the inverse gamma distribution with mean c and an infinite standard deviation.

c The data density p(YT|Mₗ) is estimated using Geweke’s (1999) modified harmonic mean estimator.

Finally, notice that Ricardian equivalence implies that lump-sum tax shocks only affect taxes and debt in the model without default risk whereas shocks to the foreign debt share only affect the division among foreign and domestic debt. As we observe no other fiscal variable other than domestic debt and government consumption, there is a stochastic singularity problem. To address this issue, we allow the innovations εₜ and ε₉ to be correlated and estimate the degree of correlation σₗ,σ₉. A positive correlation coefficient would then point towards tax-financed changes in government consumption.

5.4 Discussion of results

The discussion of results is organized as follows. In Section 5.4.1 we compare the basic model without default risk and the augmented model in terms of parameter
estimates, posterior odds, variance decompositions, and forecasting performance. In Section 5.4.2 we implement several counterfactual experiments based on the estimated model to understand the amplification channels due to default risk. We then compare the estimated impulse responses from both model versions in Section 5.4.3. Robustness checks are deferred to Section 5.4.4.

5.4.1 Model comparison

Parameters and data densities

Table 5.1 reports the posterior means of the estimated parameters, their 90% probability intervals and the log (marginal) data densities for both model versions. The following results stand out. For the model with sovereign risk ($M_1$), the inverse intertemporal substitution elasticity $\sigma$, the degree of price stickiness $\phi$, and the degree of openness $\vartheta$ are broadly in line with existing estimates for small open economies (e.g. Justiniano and Preston, 2010b; Lubik and Schorfheide, 2007), but not entirely so for the model without sovereign risk ($M_2$). A striking result is that $M_2$ has a significantly higher $\sigma$ than $M_1$. We provide an interpretation of this result below.

The estimated default elasticity $\Phi$ equals 0.29 in $M_1$, which implies that the expected default rate is highly debt-elastic. This result confirms findings of Budina and van Wijnbergen (2008) showing that higher debt service obligations of the Turkish government have led to stronger expectations that these debt obligations might not be met. Furthermore, both the inflation response in the monetary reaction function $\alpha_\pi$ and the tax feedback $\kappa$ are larger in $M_1$ than in $M_2$. In line with these results, GMM estimates of the implicit reaction function of the CBRT by Berument and Malatyali (2000) indicate a relatively large response of nominal rates to inflation when the authors control for fiscal developments.\footnote{Berument and Malatyali's (2000) results are however not directly comparable to ours due to differences in the specification of the reaction function and the estimation sample.}

Another striking result is that most of the standard deviations of the innovations are significantly larger in $M_2$, in particular the domestic demand innovations and government consumption innovations as well as the measurement errors on the real exchange rate. An exception is the standard deviation of the foreign debt share which is how-
Sovereign Default Risk and Macroeconomic Fluctuations

ever not well identified in $M_2$. Finally, a formal model comparison clearly supports the model with sovereign risk. The Bayes factor in favor of $M_1$ relative to $M_2$ is $p(Y^T|M_1)/p(Y^T|M_2) = \exp(895.51 - 748.23)$ which is equal to $9.2 \times 10^6$, indicating strong support for $M_1$ over $M_2$ conditional on the observed data.

**Default premia and effective interest rates**

Why does $M_1$ provide such a better fit to the observed data than $M_2$? To gain an intuition, notice that combining equations (5.11) and (5.19) in Appendix 5.A and using $E_t \hat{\delta}_{t+1} = \rho \hat{z}_t$ yields the following representation of the households’ consumption Euler equation:

$$\sigma(E_t \hat{c}_{t+1} - \hat{c}_t) = \hat{R}_{H,t} - E_t \hat{\pi}_{t+1} - (1 - \bar{\delta})^{-1} E_t \hat{\delta}_{t+1} - (1 - \rho) \hat{z}_t. \quad (5.10)$$

Suppose that expected consumption growth $E_t \hat{c}_{t+1} - \hat{c}_t$ shows “different” dynamics than the expected real interest rate $\hat{R}_{H,t} - E_t \hat{\pi}_{t+1}$. Indeed, according to both models, estimated consumption growth was low in the first half of the sample whereas the real interest rate was relatively high. Such dynamics could be reconciled with (5.10) in the following three ways:

1) Suppose that $E_t \hat{\delta}_{t+1} = 0$ for all $t$. With $1 - \rho > 0$, positive demand shocks $\hat{z}_t$ could make (5.10) hold if $E_t \hat{c}_{t+1} - \hat{c}_t$ is temporarily low relative to $\hat{R}_{H,t} - E_t \hat{\pi}_{t+1}$. Households would save less after a positive demand shock even if the real interest rate is high since they have a preference for temporarily higher consumption.

2) Alternatively, set both $E_t \hat{\delta}_{t+1} = 0$ and $\hat{z}_t = 0$ for all $t$. A relatively large value on the inverse intertemporal substitution elasticity $\sigma$ would increase the households’ preferences for a smooth consumption path even if the real interest rate is not smooth.

3) A positive expected default rate can balance (5.10) with relatively small demand shocks and a moderate value of $\sigma$. Households would then invest less when the real interest rate is high due to stronger default fears, and vice versa.

All three explanations are relevant to understand our estimation results. First, according to the results for $\sigma_z$ in Table 5.1, large demand shocks occur in $M_2$ whereas $M_1$ requires much smaller and less persistent shocks. Second, the estimated value of
Figure 5.1: Estimated expected default rate $E_t \tilde{\delta}_{t+1}$ and EMBIG Turkey spreads

Notes. The default rate is the estimate implied by the Kalman smoother at the posterior mean (1994Q3-2008Q2); source of EMBIG spreads (monthly data): J.P. Morgan and Bloomberg; “USD” indicates spreads on U.S. dollar Brady bonds and loans over U.S. treasury bonds (08/1998-06/2008); “Euro” indicates spreads on euro denominated bonds and loans over German bunds (05/1999-06/2008).

$\sigma$ is more than 8 times higher in $M_2$, generating a strong preference for consumption smoothing. Third, default premia were relatively high before Turkey’s financial crisis in 2000-2001 but they have declined since then as can be seen from Figure 5.1 which plots the expected default rate $E_t \tilde{\delta}_{t+1}$ from $M_1$ implied by the Kalman smoother at the posterior mean (solid line). Therefore, the effective real interest rate net of default risk $\tilde{R}_{H,t} = E_t \tilde{\pi}_{t+1} - (1 - \tilde{\delta})^{-1} E_t \tilde{\delta}_{t+1}$ shows smoother dynamics than the actual real rate, which are easier to reconcile with the estimated path of expected consumption growth.

To gauge the plausibility of the magnitude and the dynamics of the estimated expected default rate, Figure 5.1 also plots the EMBIG spreads on (i) U.S. dollar de-
nominated Turkish bonds over U.S. treasury bonds and (ii) Euro denominated Turkish bonds over German bunds.\textsuperscript{11} There is a strong co-movement, although the EMBIG indicates smaller default premia before and during the 2000-2001 crisis and somewhat larger rates afterwards. The correlations between the model-implied default rate and (i) and (ii) are 0.74 and 0.59, respectively. The default premium implied by our model thus compares well with those alternative estimates of default risk in terms of their magnitude and dynamics.

Variance decomposition

To assess the importance of alternative structural shocks in driving the expected default rate and other selected variables, Table 5.2 reports their unconditional posterior variance decomposition. Regarding the default rate ($M_1$ only), economic shocks contribute 84 percent and policy shocks 16 percent to its variation. In $M_1$, most of the variation in consumption, output, debt, the nominal interest rate, and inflation is attributed to productivity shocks. In $M_2$, in line with the large estimated value of $\sigma_z$, domestic demand shocks explain almost all of the variation in consumption whereas foreign demand shocks explain most of the fluctuations of output, the interest rate, and inflation, while most of the fluctuations in government debt are explained by lump-sum tax shocks. The presence of default beliefs therefore leads to significant changes in the importance of particular shocks in driving the dynamics of the observed variables.\textsuperscript{12}

Forecasting performance

Figure 5.2 compares selected observed variables and their one-step ahead forecasts implied by the two model versions. While both models seem to forecast output, consumption, and debt fairly well, it seems that $M_1$ generates better forecasts of inflation and the nominal interest rate. In particular, the interest rate forecasts produced by $M_2$ are excessively smooth compared to the observed data. Table 5.3 reports mean

\textsuperscript{11}All variables are reported in basis points. The estimated expected default rate ($E_t \delta_{t+1}$) was converted into levels ($E_t \delta_{t+1}$) by adding the steady state value $\delta$.

\textsuperscript{12}Shocks to the foreign debt share only affect the division of government debt among foreign and domestic debt but not its overall level. Those shocks therefore do not have any impact on the default rate and thus also do not affect the remaining variables, which explains the zeroes in the row for $\varepsilon_f$ in Table 5.2.
Table 5.2: Posterior variance decomposition\textsuperscript{a}

<table>
<thead>
<tr>
<th>Economic shocks</th>
<th>\textit{With sovereign risk (M\textsubscript{1})}</th>
<th>\textit{No sovereign risk (M\textsubscript{2})}</th>
</tr>
</thead>
<tbody>
<tr>
<td>Def. rate</td>
<td>Output</td>
<td>Consum.</td>
</tr>
<tr>
<td>Productivity $\varepsilon_a$</td>
<td>77.1</td>
<td>96.8</td>
</tr>
<tr>
<td>Dom. demand $\varepsilon_z$</td>
<td>3.6</td>
<td>1.5</td>
</tr>
<tr>
<td>For. demand $\varepsilon_c$</td>
<td>2.0</td>
<td>0.2</td>
</tr>
<tr>
<td>For. prices $\varepsilon_{\pi^*}$</td>
<td>1.0</td>
<td>0.1</td>
</tr>
<tr>
<td>\textit{Total}\textsuperscript{b}</td>
<td>83.7</td>
<td>98.6</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Policy shocks</th>
<th>\textit{With sovereign risk (M\textsubscript{1})}</th>
<th>\textit{No sovereign risk (M\textsubscript{2})}</th>
</tr>
</thead>
<tbody>
<tr>
<td>Int. rate $\varepsilon_R$</td>
<td>4.8</td>
<td>1.4</td>
</tr>
<tr>
<td>Gov. consum. $\varepsilon_g$</td>
<td>10.4</td>
<td>0.1</td>
</tr>
<tr>
<td>Lump-sum tax $\varepsilon_{\tau}$</td>
<td>1.1</td>
<td>0.0</td>
</tr>
<tr>
<td>For. debt $\varepsilon_f$</td>
<td>0.0</td>
<td>0.0</td>
</tr>
<tr>
<td>\textit{Total}\textsuperscript{b}</td>
<td>16.3</td>
<td>1.5</td>
</tr>
</tbody>
</table>

\textsuperscript{a} The table entries refer to the contribution of individual shocks to the unconditional variances (in percent) of observables at the posterior mean.

\textsuperscript{b} Some of the totals do not sum up to 100\% due to rounding errors.
Figure 5.2: Observed variables and their one-step ahead forecasts

Notes. Quarterly data, 1994:3-2008:2; one-step ahead forecasts are computed by the Kalman filter at the posterior mean; real variables are measured in percentage deviations from a linear trend, nominal variables are demeaned and in annualized percentage terms.

forecast errors (ME) and root mean squared forecast errors (RMSE) based on the one-step ahead forecasts. The RMSE are useful to judge the overall predictive performance of the two model versions. The ME help to judge whether any variable is repeatedly over- or underpredicted. The ME indicate that $M_2$ tends to overpredict the real exchange rate and underpredict government consumption. Both models tend to overpredict inflation and the nominal interest rate. Most of the RMSE are smaller for $M_1$, in particular the RMSE of the nominal interest rate, the real exchange rate, and government consumption. The RMSE of consumption is smaller in $M_2$, but recall that the
### Table 5.3: One-step ahead forecast errors

<table>
<thead>
<tr>
<th>Variable</th>
<th>Mean forecast error ME&lt;sup&gt;b&lt;/sup&gt;</th>
<th>Root mean squared forecast error RMSE&lt;sup&gt;c&lt;/sup&gt;</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>With sov. risk (M&lt;sub&gt;1&lt;/sub&gt;)</td>
<td>No sov. risk (M&lt;sub&gt;2&lt;/sub&gt;)</td>
</tr>
<tr>
<td>Output</td>
<td>-0.01</td>
<td>-0.03</td>
</tr>
<tr>
<td>Consumption</td>
<td>0.00</td>
<td>0.00</td>
</tr>
<tr>
<td>Inflation</td>
<td>0.14</td>
<td>0.11</td>
</tr>
<tr>
<td>Nom. interest rate</td>
<td>0.27</td>
<td>0.16</td>
</tr>
<tr>
<td>Domestic gov. debt</td>
<td>0.06</td>
<td>-0.05</td>
</tr>
<tr>
<td>Gov. consumption</td>
<td>0.00</td>
<td>-1.19</td>
</tr>
<tr>
<td>Real exch. rate</td>
<td>-0.05</td>
<td>0.40</td>
</tr>
<tr>
<td>For. consumption</td>
<td>0.03</td>
<td>0.02</td>
</tr>
<tr>
<td>For. inflation</td>
<td>-0.01</td>
<td>0.01</td>
</tr>
</tbody>
</table>

<sup>a</sup> The forecast errors $F_t$ are computed as the difference between the observed variables $Y_t$ and their one-step ahead forecasts $Y^f_t$ as $F_t = Y_t - Y^f_t$, where $Y_t$ and $Y^f_t$ are measured in percent.

<sup>b</sup> The mean forecast errors are computed as $\text{ME} = \frac{1}{T-1} \sum_{t=1}^{T} F_t$.

<sup>c</sup> The mean squared forecast errors are computed as $\text{RMSE} = \left( \frac{1}{T-1} \sum_{t=1}^{T} F_t^2 \right)^{1/2}$.

The bulk of consumption fluctuations is driven by domestic demand shocks in that model, indicating empirical shortcomings of the basic consumption Euler equation. Overall, the model with sovereign risk thus performs better in terms of forecasting performance.

#### 5.4.2 Amplification channels

We now describe several counterfactual experiments to isolate individual features of the model and to examine their importance in amplifying shocks. All of the following experiments are based on the estimated model with sovereign risk, which is referred to as the benchmark model: (i) the default elasticity $\Phi$ is set to zero, (ii) the degree of openness $\vartheta$ is set to zero, (iii) the foreign debt share $\bar{f}$ is set to zero, and (iv) the inverse intertemporal substitution elasticity $\sigma$ is set to 15.98, its posterior mean in the model without sovereign risk. Figure 5.3 shows the impulse responses to a negative one percent productivity shock for the four experiments.

When the default elasticity $\Phi$ is set to zero (thick dashed line), Ricardian equivalence holds. The negative productivity shock causes a decline in output and a rise in intermediate goods firms’ marginal costs, leading to an increase in prices and an ap-
Figure 5.3: Estimated and counterfactual impulse responses to a productivity shock.

Notes. Based on model with sovereign risk; productivity shock is normalized to minus one percent; estimated impulse responses are calculated at the posterior mean and counterfactual impulse responses are calculated by changing one parameter at a time; real variables are measured in percentage deviations from steady state, nominal variables in absolute (annual) percentage point deviations from steady state.

preciation of the real exchange rate. Domestic consumption falls due to international risk sharing and expenditure switching of domestic and foreign households. Domestic output therefore declines further. The monetary authority reacts to higher inflation by increasing the nominal interest rate. Government debt falls initially, due to the immediate beneficial exchange rate effect on foreign debt, but it rises afterwards due to higher debt service obligations resulting from higher interest rates.

In the benchmark model with a positive default elasticity (solid line) the real value of debt affects the expected effective rate of return to investors, which alters the dynamics through various channels. As before, higher inflation leads to higher nominal
interest rates, higher debt service obligations, and higher debt. However, through the negative feedback from debt on its expected return savings tend to be lower and current consumption tends to be higher, leading to inflationary pressures. In order to contain inflation, the monetary authority raises the nominal interest rate by more than in experiment (i). Higher nominal rates in turn imply higher debt servicing costs and debt levels, further increasing expected default rates. The latter leads to additional inflationary pressures. The initial increase of inflation is thus amplified through the presence of default beliefs.

The exchange rate channel further enhances the inflationary pressures. The tendency of domestic current consumption to rise, resulting from the negative feedback from government debt on its return, feeds into pressure on the real exchange rate to depreciate through international risk sharing. A real depreciation would lead to expenditure switching of domestic households and increasing demand of foreign households for home goods. Moreover, domestic households would demand a higher nominal wage since the price level of aggregate consumption would rise due to higher prices of imported goods. To counteract these inflationary pressures, the monetary authority needs to raise the nominal interest rate even more. Conversely, in a closed economy (dashed-dotted line) the impact of the productivity shock on inflation and the nominal interest rate is significantly muted, such that the responses of government debt and the expected default rate are also smaller.

Similarly, without foreign currency denominated debt (solid line with dots) the effects on inflation, the nominal interest rate, and debt are smaller. In the absence of foreign debt, nominal depreciation does not trigger additional default beliefs due to debt revaluation. Moreover, the devaluing effect of higher domestic inflation on the stock of real debt is more pronounced if debt is only denominated in domestic currency. Finally, for high values of the inverse intertemporal substitution elasticity $\sigma$ (bars) the response of consumption to an increase of the nominal interest rate is substantially muted since households have a strong preference for a smooth consumption path. The effectiveness of high nominal rates in containing inflationary pressures is reduced such that higher nominal rates are required. Higher nominal rates in turn imply higher debt and, again, higher expected default rates.
Figure 5.4: Estimated impulse responses to a productivity shock in both models

Notes. Productivity shock is normalized to minus one percent with persistence as estimated in the model with sovereign risk; impulse responses are calculated at the posterior mean; real variables are measured in percentage deviations from steady state, nominal variables in absolute (annual) percentage point deviations from steady state.

5.4.3 Estimated impulse responses

Having analyzed particular model features in isolation, we now compare the estimated impulse responses implied by the models with default risk \((M_1)\) and without default risk \((M_2)\). We focus again on the responses to a negative productivity shock which is normalized to have the persistence from \(M_1\), i.e. we set \(\rho_a = 0.96\). The dashed lines in Figure 5.4 show the impulse responses implied by \(M_2\). The negative productivity shock causes a rise in inflation and a real appreciation. Domestic consumption and output fall. The monetary authority increases the nominal interest rate, government debt falls initially, and then shows a persistent increase due to higher debt service obligations.
resulting from higher interest rates.

In view of the counterfactuals above, the amplification of the responses of inflation, the nominal interest rate, and domestic debt implied by $M_1$ (solid line) can mainly be attributed to the presence of default beliefs. The different responses of output, consumption, and the real exchange rate seem to be mainly driven by the lower estimated substitution elasticity $\sigma$, which implies a more pronounced response of consumption to movements in the real effective interest rate. However, the effect on the real exchange rate is muted, since with a low $\sigma$ fluctuations in domestic consumption only feed into small variations of the real exchange rate via international risk sharing. Finally, the higher estimated degree of openness further enhances the effects of default risk on nominal variables, government debt, and consumption.

5.4.4 Sensitivity checks

As a final step of the analysis, alternative versions of the benchmark model with default risk are estimated to check the sensitivity of the estimation results. The results are summarized in Table 5.4.

First, we introduce output and exchange rate stabilization terms in the monetary authority’s reaction function to check whether the CBRT was targeting output and to capture the fact that before 2001 the CBRT’s monetary policy strategy included nominal exchange rate targeting (see Gormez and Imaz, 2007). Thus (5.3) is replaced by the modified reaction function

$$R_{H,t} / R_H = (\pi_t / \pi)^{\alpha_\pi} (y_t / y)^{\alpha_y} (X_t / X_{t-1})^{\alpha_X} \exp(\varepsilon_{R,t}),$$

where $X_t / X_{t-1}$ is the rate of nominal depreciation. We use uniform priors on the range $[-10, 10]$ for both the output feedback $\alpha_y$ and the exchange rate feedback $\alpha_X$. While the estimated exchange rate feedback is fairly large (0.44), confirming that exchange rate stabilization was a concern of the CBRT, the estimated output feedback is close to zero (-0.01). The size of the inflation feedback $\alpha_\pi$ decreases compared to the benchmark model from 1.83 to 1.43, but the remaining parameter estimates do not change significantly. The data density falls, which suggests that the additional feedback terms
do not improve the overall fit of the model.

Second, we check whether it matters if we drop the exogenous persistence mechanism in the domestic households’ consumption Euler equation due to persistent demand shocks and add an endogenous persistence mechanism instead. In particular, we incorporate external habit formation in consumption, as in Adolphson et al. (2007) or Justiniano and Preston (2010a), and set $\rho_z = 0$ when estimating the model. The (domestic and foreign) households’ preferences are modified accordingly:

$$E_0 \sum_{t=0}^{\infty} \beta^t \left[ \exp(\varepsilon_{zt}) \frac{1}{1-\sigma} (c_t - h\tilde{c}_{t-1})^{1-\sigma} - \frac{1}{1+\eta} r_t^{1+\eta} \right],$$

where $h \in (0,1)$ and $\tilde{c}_{t-1}$ denotes aggregate domestic consumption, which is taken as

### Table 5.4: Sensitivity of parameter estimates and data densities\(^a\)**

<table>
<thead>
<tr>
<th>Parameter</th>
<th>Prior(^b)</th>
<th>Benchmark model</th>
<th>Mod. mon. rule</th>
<th>Habit formation</th>
<th>Price indexation</th>
<th>Smaller meas. error(^b)</th>
<th>Max. likel. estim.</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\sigma$</td>
<td>U(0,20)</td>
<td>1.87</td>
<td>1.80</td>
<td>1.20</td>
<td>1.68</td>
<td>0.57</td>
<td>2.03</td>
</tr>
<tr>
<td>$h$</td>
<td>U(0,1)</td>
<td>—</td>
<td>—</td>
<td>0.58</td>
<td>—</td>
<td>—</td>
<td>—</td>
</tr>
<tr>
<td>$\phi$</td>
<td>U(0,1)</td>
<td>0.46</td>
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<td>0.51</td>
<td>0.39</td>
<td>0.17</td>
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<tr>
<td>$\nu$</td>
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<td>—</td>
<td>0.47</td>
<td>—</td>
<td>—</td>
<td>—</td>
</tr>
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<td>$\vartheta$</td>
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<td>0.13</td>
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</tr>
<tr>
<td>$\Phi$</td>
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<td>0.29</td>
<td>0.26</td>
<td>0.31</td>
<td>0.28</td>
<td>0.31</td>
</tr>
<tr>
<td>$\alpha_x$</td>
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<td>1.43</td>
<td>1.81</td>
<td>1.87</td>
<td>2.14</td>
<td>1.77</td>
</tr>
<tr>
<td>$\alpha_y$</td>
<td>U(-10,10)</td>
<td>—</td>
<td>—</td>
<td>—</td>
<td>—</td>
<td>—</td>
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<td>0.44</td>
<td>0.48</td>
<td>0.46</td>
<td>0.58</td>
<td>0.40</td>
</tr>
<tr>
<td>$\alpha_t$</td>
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<td>0.62</td>
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<tr>
<td>$\rho_z$</td>
<td>U(0,1)</td>
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<td>0.74</td>
<td>—</td>
<td>0.72</td>
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<td>0.95</td>
<td>0.97</td>
<td>0.94</td>
<td>0.92</td>
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</tr>
<tr>
<td>$\rho_b$</td>
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<td>0.44</td>
<td>0.48</td>
<td>0.46</td>
<td>0.58</td>
<td>0.40</td>
</tr>
<tr>
<td>$\rho_f$</td>
<td>U(0,1)</td>
<td>0.90</td>
<td>0.89</td>
<td>0.90</td>
<td>0.89</td>
<td>0.87</td>
<td>0.90</td>
</tr>
<tr>
<td>$\sigma_z$</td>
<td>IG(0.05)</td>
<td>0.05</td>
<td>0.05</td>
<td>0.08</td>
<td>0.05</td>
<td>0.01</td>
<td>0.06</td>
</tr>
<tr>
<td>$\sigma_a$</td>
<td>IG(0.05)</td>
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<td>0.03</td>
<td>0.03</td>
<td>0.03</td>
<td>0.02</td>
<td>0.03</td>
</tr>
<tr>
<td>$\sigma_g$</td>
<td>IG(0.05)</td>
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<td>0.04</td>
<td>0.04</td>
<td>0.04</td>
<td>0.04</td>
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</tr>
<tr>
<td>$\sigma_f$</td>
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<td>0.25</td>
<td>0.25</td>
<td>0.26</td>
<td>0.30</td>
<td>0.24</td>
</tr>
<tr>
<td>$\sigma_r$</td>
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<td>0.07</td>
<td>0.07</td>
<td>0.08</td>
<td>0.09</td>
<td>0.07</td>
</tr>
<tr>
<td>$\sigma_R$</td>
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<td>0.07</td>
<td>0.06</td>
<td>0.06</td>
<td>0.07</td>
<td>0.06</td>
</tr>
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<td>0.11</td>
<td>0.11</td>
<td>0.11</td>
<td>—</td>
<td>0.11</td>
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<tr>
<td>$\rho_{z,g}$</td>
<td>U(-1,1)</td>
<td>0.79</td>
<td>0.79</td>
<td>0.74</td>
<td>0.78</td>
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<td>0.90</td>
</tr>
<tr>
<td>Log data density</td>
<td></td>
<td>895.51</td>
<td>889.90</td>
<td>891.19</td>
<td>895.91</td>
<td>862.77</td>
<td>—</td>
</tr>
</tbody>
</table>

\(^a\) See Table 5.1 for details on the estimation and the parameterization of the prior distributions.

\(^b\) For the specification with smaller measurement errors, the std. deviation $\sigma_q$ is calibrated to 0.05.
given by the individual households. We elicit a U(0, 1) prior on \( h \) in consistence with its theoretical domain. The first-order conditions for (domestic and foreign) consumption become 
\[
\lambda_t = \exp(\varepsilon_{z,t}) (c_t - hc_{t-1})^{-\sigma} \quad \text{and} \quad \lambda^*_t = (c^*_t - hc^*_{t-1})^{-\sigma},
\]
where the equilibrium conditions \( c_t = \bar{c}_t \) and \( c^*_t = \bar{c}^*_t \) have been imposed for all \( t \). Table 5.4 shows that, although \( h \) is fairly large with an estimated value of 0.58, the marginal data density falls and most parameters do not change significantly. The only exceptions are the substitution elasticity \( \sigma \), which decreases from 1.87 to 1.20, and the volatility of demand shocks \( \sigma_z \), which increases from five percent to eight percent.

Third, to check whether incorporating an endogenous persistence mechanism for inflation affects our results, we allow for partial indexation to past and steady state producer price inflation following Smets and Wouters (2007). The firms of fraction \( \phi \) that do not set prices optimally in period \( t \) are thus assumed to adjust prices according to the indexation rule 
\[
P^i_{H,t} = \bar{P}_{H,t-1}(\pi_{H,t-1}^{-1})^{},
\]
where \( \iota \in [0, 1] \) measures the degree of indexation to past producer price inflation. For \( \iota = 0 \) prices are fully indexed to the steady state inflation rate, as in the benchmark model, while for \( \iota = 1 \) prices are fully indexed to the previous period’s inflation rate. Partial indexation leads to a modified log-linearized Phillips curve (5.16):
\[
\dot{\pi}_{H,t} = (1 - \phi)(1 - \phi \beta)[\phi(1 + \iota \beta)]^{-1} \hat{m}c_t + \iota(1 + \iota \beta)^{-1} \pi_{H,t-1} + \beta(1 + \iota \beta)^{-1} E_t \hat{\pi}_{H,t+1}.
\]

We elicit a U(0, 1) prior for \( \iota \) in consistence with its theoretical domain. The estimated \( \iota \) is positive but less than half (0.47) and the degree of price stickiness \( \phi \) falls from 0.46 to 0.39, but the remaining estimates are stable. The data density increases by a few decimal points only. These results are in line with findings of Celasun, Gelos, and Prati (2004) indicating that the degree of inflation persistence induced by price indexation is comparably low in Turkey. Note that the estimated default elasticity \( \Phi \) increases slightly such that default fears remain a relevant concern when controlling for backward-looking behavior in price setting.

Fourth, to gauge the importance of the measurement error on the real effective exchange rate, we calibrate its standard deviation \( \sigma_q \) to 0.05, i.e. its prior mean. Several estimates change, in particular the values of \( \sigma, \phi, \theta, \) and \( \alpha_\pi \), but the default
elasticity does not change significantly. Finally, we estimate the benchmark model by constrained maximum likelihood (ML), restricting the model parameters on their theoretically feasible range according to the domains in Table 5.1. The ML results are broadly similar to the Bayesian estimation results. The default elasticity is again highly debt-elastic with an estimated value of 0.31.

The estimated expected default rates from all estimated versions of the model are plotted in Figure 5.5. The results are very similar across models, the only exception being the model where $\sigma_q$ is not estimated. Even in this case the estimated default rate does however show a strong correlation with the estimate from the benchmark model. Overall, the estimation results thus seem robust to various changes in the specification of the model and its empirical implementation.
5.5 Conclusion

This chapter has set up an empirical DSGE model of a small open economy where a perceived risk of sovereign debt default leads to a time-varying default premium on government bonds in domestic and foreign households’ consumption Euler equation. The default premium is linked to the stock of total government debt. We use the model to match Turkish time series data showing remarkable fluctuations in inflation, nominal interest rates, and government debt.

Our results show that the introduction of sovereign default risk helps to explain business cycles in Turkey in important ways. The presence of default beliefs not only amplifies the fluctuations of nominal variables and government debt, and thus helps to explain the high volatility of these variables, but also introduces a mutual link between the persistence of nominal variables and the persistence of government debt. Key to understand the improved fit is the consumption Euler equation for investments in government bonds. The time-varying default premium helps to reconcile the observed interest rate movements with expected consumption growth and inflation by reducing the effective rate of return on investing in government bonds.

To sum up, this chapter has shown that it is possible to build a realistic NOEM model to explain macroeconomic fluctuations in an emerging market economy where sovereign default risk is a relevant concern. We conclude by noting that there is empirical evidence that the relationship between government debt and default premia also contains non-linear elements (e.g. Bayoumi, Goldstein, and Woglom, 1995), such that the estimation approach used in this paper may only provide an incomplete picture of the link between sovereign risk and business cycle dynamics. Implementing empirical DSGE models with fiscal sustainability concerns by non-linear estimation procedures would therefore be a useful direction for future research.

5.A Equilibrium conditions

This appendix contains the extensive representation of the symmetric equilibrium conditions. The log deviation and absolute deviation of a variable $x_t$ from its non-stochastic steady state value $x$ are denoted by $\hat{x}_t$ and $\tilde{x}_t$, respectively. Variables with bars denote
steady state values that are taken as given.

**Households:**

\[
\hat{\lambda}_t = \hat{z}_t - \sigma \hat{c}_t, \quad (5.11)
\]
\[
\eta \hat{m}_t = \hat{\lambda}_t + \hat{w}_t, \quad (5.12)
\]
\[
\hat{\lambda}^*_t = -\sigma \hat{c}^*_t. \quad (5.13)
\]

**Production and pricing:**

\[
\hat{y}_{H,t} = \hat{a}_t + \hat{m}_t, \quad (5.14)
\]
\[
\hat{m}c_t = \frac{\vartheta}{1 - \vartheta} \hat{q}_t + \hat{w}_t - \hat{a}_t, \quad (5.15)
\]
\[
\hat{\pi}_{H,t} = (1 - \phi) (1 - \phi \beta) \frac{\phi}{\vartheta} \hat{m}c_t + \beta E_t \hat{\pi}_{H,t+1}, \quad (5.16)
\]
\[
\hat{\pi}_t = \hat{\pi}_{H,t} + \vartheta (\hat{q}_t - \hat{q}_{t-1}). \quad (5.17)
\]

**Capital market:**

\[
\hat{\lambda}^*_t = \hat{q}_t + \hat{\lambda}_t, \quad (5.18)
\]
\[
\hat{\lambda}_t = E_t \hat{\lambda}_{t+1} + R_{H,t} - E_t \hat{\pi}_{t+1} - \frac{1}{1 - \delta} E_t \hat{\delta}_{t+1}, \quad (5.19)
\]
\[
\hat{\lambda}^*_t = E_t \hat{\lambda}^*_{t+1} + \hat{R}_{F,t} - E_t \hat{\pi}^*_{t+1} - \frac{1}{1 - \delta} E_t \hat{\delta}_{t+1}, \quad (5.20)
\]
\[
E_t \hat{\delta}_{t+1} = \Phi (1 - \delta) (1 + \bar{f}) E_t \hat{b}_{t+1}. \quad (5.21)
\]

**Policy:**

\[
\hat{q}_t + b_{F,t} - \hat{R}_{F,t} = \hat{f}_t + \hat{b}_{H,t} - \hat{R}_{H,t}, \quad (5.22)
\]
\[
(1 + \bar{f}) \hat{b}_t = \hat{b}_{H,t-1} - \hat{\pi}_t + \bar{f} \left( \hat{q}_t + b_{F,t-1} - \hat{\pi}^*_t \right), \quad (5.23)
\]
\[
\hat{b}_{H,t} - \hat{R}_{H,t} + \bar{f} \left( \hat{q}_t + b_{F,t} - \hat{R}_{F,t} \right) = \frac{\kappa + \beta (1 - \delta) - 1}{\beta (1 - \delta) (1 + \bar{f})^{-1}} \left( \hat{g}_t - \frac{\vartheta}{1 - \vartheta} \hat{q}_t \right) + \frac{(1 - \kappa)(1 + \bar{f})}{\beta (1 - \delta)} \hat{b}_t - \varepsilon_{r,t} \quad (5.24)
\]
\[
\hat{R}_{H,t} = \alpha \hat{\pi}_t + \varepsilon_{R,t}. \quad (5.25)
\]
Market clearing:

\[
\hat{y}_{H,t} = (1 - \theta) \hat{s}_c \hat{c}_t + [1 - (1 - \theta) \hat{s}_c - \bar{s}_g] \hat{c}_t^*
\]

\[
+ \left( \frac{\hat{s}_c - \bar{s}_g}{1 - \theta} \right) \hat{q}_t + \bar{s}_g \hat{g}_t.
\]  

(5.26)

Stochastic processes:

\[
\hat{z}_t = \rho_z \hat{z}_{t-1} + \varepsilon_{z,t},
\]  

(5.27)

\[
\hat{a}_t = \rho_a \hat{a}_{t-1} + \varepsilon_{a,t},
\]  

(5.28)

\[
\hat{g}_t = \rho_g \hat{g}_{t-1} + \varepsilon_{g,t},
\]  

(5.29)

\[
\hat{f}_t = \rho_f \hat{f}_{t-1} + \varepsilon_{f,t},
\]  

(5.30)

\[
\rho_{\alpha_0,\zeta}^c = \rho_{\alpha_1,\zeta}^c \hat{c}_{t-1} + \rho_{\alpha_2,\zeta}^c \hat{c}_{t-2} + \rho_{\alpha_3,\zeta}^c \hat{c}_{t-3} + \rho_{\alpha_4,\zeta}^c \hat{c}_{t-4} + \varepsilon_{c,t},
\]  

(5.31)

\[
\rho_{\alpha_0,\zeta}^g = \rho_{\alpha_1,\zeta}^g \hat{g}_{t-1} + \rho_{\alpha_2,\zeta}^g \hat{g}_{t-2} + \rho_{\alpha_3,\zeta}^g \hat{g}_{t-3} + \rho_{\alpha_4,\zeta}^g \hat{g}_{t-4} + \varepsilon_{g,t},
\]  

(5.32)

where \( \bar{s}_c \) and \( \bar{s}_g \) denote the ratios of private consumption and government consumption over GDP, respectively.

The rational expectations equilibrium of this model is then the set of sequences

\[
\{ \hat{c}_t, \hat{c}_t^*, \hat{\lambda}_t, \hat{\lambda}_t^*, \hat{z}_t, \hat{n}_t, \hat{w}_t, \hat{a}_t, \hat{y}_{H,t}, \hat{m}_{c,t}, \hat{q}_t, \hat{\pi}_{H,t}, \hat{\pi}_t, \hat{\pi}_t^*, \hat{b}_t, \hat{b}_{H,t}, \hat{f}_t, \hat{g}_t, \hat{R}_{H,t}, \hat{R}_{F,t}, \hat{\delta}_t \}_{t=0}^{\infty}
\]

satisfying (5.11)-(5.32) and the transversality conditions, for given initial asset endowments \( B_{H,-1}, B_{F,-1} \), and \( S_0 \) and initial price levels \( P_{H,-1} \) and \( P_{F,-1} \). The i.i.d. innovations are given by \( \{ \varepsilon_{z,t}, \varepsilon_{a,t}, \varepsilon_{f,t}, \varepsilon_{g,t}, \varepsilon_{R,t}, \varepsilon_{\pi,t}, \varepsilon_{c,t}, \varepsilon_{g,t} \} \) for \( t=0 \).

5.B Detailed data description

This appendix provides details on data sources, definitions and the construction of foreign variables. The data is seasonally adjusted and the consumer price index is used to construct real variables with base year 1998, if they are only available in nominal.
terms from the original source.

The domestic variable definitions and their sources are as follows:

- \( \text{GDP}_t \): Real gross domestic product, CBRT.
- \( \text{CONS}_t \): Real private consumption expenditure, CBRT.
- \( \text{GOV}_t \): Real government consumption expenditure, CBRT.
- \( \text{DEBT}_t \): Domestic debt position of the treasury, CBRT.
- \( \text{INT}_t \): Annual net interest rate for 3-month treasury bills, constructed from data obtained from the CBRT; if 3-month bills were not issued in some quarter, we use the closest maturity available.
- \( \text{INF}_t \): Annualized rate of change of the quarterly CPI, State Institute of Statistics Turkey.
- \( \text{REER}_t \): Real CPI-based effective exchange rate, OECD main economic indicators.

The foreign variables are constructed from euro area real private consumption and the annual inflation rate according to the HICP index obtained from the AWM database (Fagan, Henry, and Mestre, 2005), and real U.S. personal consumption and the CPI-based U.S. inflation rate (all urban sample, all items) obtained from the BEA. Aggregate foreign consumption \( \text{CONS}^*_t \) and foreign inflation \( \text{INF}^*_t \) are computed according to the trade weights in the basket targeted by the CBRT during the exchange rate targeting period (see Gormez and lmaz, 2007). That is, the euro area obtains a weight of 0.77 and the U.S. obtains a weight of one.
Chapter 6

Summary and Conclusion

This thesis has studied the relationship between fiscal policy and the business cycle. In particular, based on a combination of empirical macroeconometric techniques and macroeconomic theory, the thesis has analyzed the impact of government expenditure policies, public debt, and sovereign default risk on macroeconomic fluctuations. The aim of this analysis has been to contribute to fill some of the gaps in contemporary macroeconomic research on fiscal policy that have become obvious in the context of the turbulent economic events over the past few years.

Chapter 2 has estimated time-varying parameters time series models for the euro area to examine changes in the macroeconomic effects of government spending. It has also analyzed the driving forces of those changes using regression analysis.

Chapter 3 has demonstrated how expectations on future government spending can be used to address the problems of econometric inference on the effects of spending changes that are caused by policy foresight by economic agents. The proposed expectations-based approach has been applied to estimate the effects of surprise changes in government spending in the United States.

Chapter 4 has assessed the effectiveness of government policies during a financial crisis in a structural macroeconomic model with frictions in financial intermediation. The model takes into account that a large fraction of government debt is typically held by financial institutions instead of directly by households.

Chapter 5 has analyzed the role of sovereign default risk in macroeconomic fluctuations in an emerging market economy based on an estimated quantitative business
cycle model for Turkey.

The results from Chapters 2 and 3 suggest that, on average, expenditure policies falling on government goods purchases are not very effective tools for short-run stabilization. Moreover and especially over time, the medium-term effects of such policies point towards significant crowding-out effects of fiscal expansions. The results of Chapter 4 suggest that, if the macroeconomic context calls for fiscal policy actions, they should be chosen carefully with a view of their financing. This chapter shows that deficit financing through a distressed financial sector can have undesirable consequences for the effectiveness of government policies. The results of Chapter 5, on the other hand, show that the observed dynamics of government debt and monetary policy behavior in interaction with sovereign default premia can explain important aspects of past business cycles in Turkey. The latter indicates that fiscal variables matter for macroeconomic fluctuations when agents view fiscal developments as potentially unsustainable even if fiscal policy remains rather passive.

The overall message that emerges from the analysis in Chapters 2 to 5 of this thesis thus calls for a prudent approach towards the use of fiscal policy, especially government expenditure policies, over the business cycle. The results of the analysis do however not necessarily support total fiscal passivism, but they do the view that fiscal measures are not very effective if they are implemented without taking into account the specific macroeconomic context. At the same time, the results suggest that an important contribution that governments can make to achieve macroeconomic stability is to ensure that fiscal variables follow a sustainable path.

The thesis has not studied a number of other relevant aspects. In particular, the role of tax and transfer policies has not been analyzed, mainly due to significant problems in the empirical measurement of autonomous changes in taxes and transfers (Chapters 2 and 3) but also a relative lack of data (Chapter 3). In addition, the models in Chapters 4 and 5 have not accounted for distortionary taxation. The main transmission channels in those models are however unlikely to be affected by an extension in this direction. Moreover, one could have looked at a broader set of government policies in Chapter 4. The specific choice of policies seems however irrelevant for the characterization and the understanding of the key mechanisms highlighted in this chapter.
On the other hand, while Chapter 5 represents a further step towards an analysis of sovereign default risk in an estimated quantitative macroeconomic model, ongoing concerns related to the exposure of financial institutions to risky government debt suggest that it may be useful to combine the framework in Chapter 4 with the relevant aspects of Chapter 5 to be able to evaluate monetary and fiscal policies in this type of environment. The latter is left for future research.
Fiscaal beleid en de conjunctuurcyclus
Het effect van overheidsuitgaven, overheidsschuld en wanbetalingsrisico bij overheden op macro-economische fluctuaties

Dit proefschrift onderzoekt de relatie tussen fiscaal-economisch beleid en conjuncturele fluctuaties. De specifieke focus ligt op de macro-economische effecten van de overheidsuitgaven, de overheidsschuld en het risico op wanbetaling door overheden. De analyse is gebaseerd op een combinatie van empirische macro-econometrische methoden en macro-economische theorie. Het doel van de analyse is om bij te dragen de hieronder vermelde lacunes in het hedendaags macro-economisch onderzoek te vullen. De relevantie van deze aspecten en het gebrek aan onderzoek op dit terrein werden gedurende de turbulente economische gebeurtenissen van de afgelopen jaren zeer duidelijk.

1. Binnen het empirisch macro-economisch onderzoek is nog steeds geen voldoende consensus over de algemene economische effecten van fiscale maatregelen. De belangrijkste redenen voor het gebrek aan consensus zijn ten eerste problemen bij het onderscheiden van discretionaire fiscale maatregelen en automatische reacties van fiscale variabelen zoals inkomensoverdrachten en belastinginkomsten op cyclische fluctuaties, ten tweede de mogelijkheid dat discretionaire maatregelen al voor hun invoering bij marktdeelnemers bekend kunnen zijn, bijvoorbeeld in het geval van aangekondigde uitgavenprogramma’s, en ten derde instabiliteit van empirische resultaten over verschillende periodes, bijvoorbeeld als gevolg van structurele economische veranderingen. De eerste twee problemen vormen een belangrijke uitdaging voor vaak gebruikte empirische methoden omdat zij schattingsresultaten die op deze methoden gebaseerd zijn sterk kunnen vertekenen,
terwijl de invloed van structurele veranderingen de interpretatie van schattingsresultaten voor overlappende perioden bemoeilijkt. Een voorbeeld van een relevante structurele verandering is de toegenomen wereldwijde economische integratie, die tot meer handel en in enkele landen wellicht tot een grotere importcomponent van een budgettaire expansie geleid kan hebben. Dit grotere importaandeel kan invloed hebben op de effectiviteit van een budgettaire expansie voor de binnenlandse economische activiteit.

2. Ondanks een toename in relevant onderzoek in de afgelopen jaren kan de macro-economische literatuur slechts relatief weinig bijdragen aan de aanbeveling van passende fiscale maatregelen tijdens een crisis met zijn oorsprong in de financiële sector zoals de recente economische en financiële crisis. Dit betreft met name de empirische literatuur, vanwege een gebrek aan gegevens over relevante crisis episodes en vergelijkbare econometrische problemen zoals onder punt 1 werden genoemd. Het gebrek aan beleidsrelvante aanbevelingen kan ook worden toegeschreven aan het feit dat relevante verbindingen tussen de reële economie en de financiële sector in de afgelopen decennia in het theoretisch economisch onderzoek verwaarloosd werden, mogelijk als gevolg van een vals gevoel van stabiliteit van de macro-economische ontwikkelingen in de grote ontwikkelde landen tot aan de recente crisis. Relevant kwesties zoals de impact van een abrupte stijging van het overheidstekort op de balansstructuur van financiële instellingen werden voor zover de auteur bekend is nog niet eerder bestudeerd in macro-economische modellen. Dergelijke vragen hebben wellicht een belangrijke rol gespeeld in de economisch-politieke discussies gedurende de recente crisis.

3. De literatuur laat ook ruimte voor verder onderzoek naar kwantitatieve economische modellen die rekening houden met de mogelijkheid van wanbetalings door nationale overheden. Het zou nuttig zijn om deze mogelijkheid met behulp van theoretische modellen te onderzoeken. Dit zou kunnen helpen het verband tussen de overheidsschuld, betalingsrisico’s en conjuncturele fluctuaties beter te begrijpen, om passende beleidsmaatregelen te kunnen aanbevelen in een omgeving waar een hoge overheidsschuld een probleem is. Slechts enkele studies hebben
de effecten van het risico van wanbetalings door overheden in structurele macroeconomic modellen onderzocht, in het bijzonder structurele modellen die met behulp van empirische methoden kunnen worden geschat die de dynamiek van de bijbehorende economische tijdreeksen volledig meeneemt. Dit soort analyse kan wellicht bijdragen aan de theoretisch-empirische beschrijving van de relevante macro-economische mechanismen.

In het licht van deze aspecten is de analyse in dit proefschrift als volgt opgebouwd. Hoofdstuk 2, gebaseerd op gezamenlijk werk met Jacopo Cimadomo en Sebastian Hauptmeier, richt zich op het probleem van instabiliteit in de macro-economische effecten van overheidsuitgaven over de tijd. In het bijzonder worden in dit hoofdstuk tijdreeks modellen met parameters die over tijd fluctueren op basis van Euro-zone data voor de periode 1980-2008 geschat om veranderingen in de effecten van overheidsconsumptie en overheidsinvesteringen in Europese landen te onderzoeken. Het hoofdstuk analyseert ook de oorzaken van deze veranderingen aan de hand van regressie analyses. De resultaten tonen aan dat het effect van overheidsuitgaven op de economische activiteit tussen 1980 en 2008 kleiner is geworden. Deze trend kan vooral worden toegeschreven aan betere beschikbaarheid van leningen en toenemende overheidsschuld. Betere beschikbaarheid van leningen kan het bijvoorbeeld voor marktdeelnemers makkelijker maken om de opbrengst van een budgettaire expansie te sparen voor hogere toekomstige belastingen. Aan de andere kant kan een hogere overheidsschuld bijvoorbeeld invloed hebben op de effectiviteit van een budgettaire expansie vanwege extra budgettaire tekorten.

Hoofdstuk 3 laat zien hoe gegevens over verwachtingen over toekomstige overheidsuitgaven kunnen worden gebruikt om de problemen van structurele vector autoregressieve econometrische technieken te verminderen, die worden veroorzaakt door het feit dat er vaakop discretionaire beleidsmaatregelen kan worden geanticipeerd door marktdeelnemers. Dit fenomeen is een grote uitdaging voor empirisch onderzoek, omdat de relevante macro-economische dynamiek in een dergelijke situatie niet door vector autoregressies op basis van actuele en historische gegevens over fiscale variabelen kan worden beschreven. Empirische resultaten kunnen daardoor onbetrouwbaar zijn als het bovenstaande fenomeen wordt genegeerd. Om dit probleem te verhelpen hebben enkele
recente studies gesuggereerd “forward-looking” variabelen in vector autoregressies te gebruiken. Het doel van dit hoofdstuk is een theoretische basis voor dit soort oplossingsrichtingen te creëren. In het bijzonder schetst het hoofdstuk een empirische benadering op basis van gegevens over verwachtingen en past deze benadering dan op Amerikaanse data toe. De empirische resultaten laten zien dat budgettaire expansies op basis van hogere overheidsuitgaven minder effectief zijn dan de conventionele vector autoregressieve methode aangeeft, waarin niet rekening wordt gehouden met anticiperend gedrag wat betreft toekomstige overheidsuitgaven.

Hoofdstuk 4, gebaseerd op gezamenlijk werk met Sweder van Wijnbergen, vergeleijkt de effectiviteit van verschillende fiscale maatregelen tijdens een financiële crisis in een structurele macro-economische model met fricties in de financiële sector. Om te komen tot een realistische beschrijving van fiscaal-financiële aspecten houdt het model rekening met het feit dat een groot deel van de bestaande staatsobligaties in handen van financiële instellingen is en niet rechtstreeks in handen van huishoudens. De maatregelen die worden overwogen omvatten zowel conjuncturele maatregelen (met name op uitgaven gebaseerde economische stimuleringsmaatregelen) als maatregelen ten gunste van de financiële sector (met name leningen en kapitaaloverdrachten). De resultaten van deze analyse tonen aan dat fiscale maatregelen ook destabiliserende effecten tijdens een crisis kunnen hebben, afhankelijk van het deel van de overheidstekort dat wordt gefinancierd via de financiële sector. De reden is dat door schulden gefinancierde maatregelen de kredietverlening aan de particuliere sector kunnen verdringen.

Hoofdstuk 5, gebaseerd op gezamenlijk werk met Malte Rieth, analyseert de invloed van het risico op wanbetalingsrisico’s bij overheden op cyclische fluctuaties aan de hand van een voorbeeld opkomend land. De analyse baseerd op een geschat kwantitatief model voor Turkije. Het hoofdstuk laat zien dat de waargenomen dynamiek van de Turkse overheidsschuld en het gedrag van het Turkse monetair beleid in interactie met de wanbetalingskans op de overheidsschuld belangrijke aspecten van in het verleden geobserveerde conjunctuurcycli in Turkije kunnen verklaren. De resultaten tonen in het bijzonder aan dat fiscale variabelen ook invloed kunnen hebben op conjuncturele fluctuaties (in overeenstemming met dit hoofdstuk over de effecten van de wanbetalingskans op de overheidsschuld op de consumptie en het opslaan van parti-
kuliere beleggers) als het fiscale beleid relatief passief is. Een mogelijke conclusie is dat een beleid waarbij overheidsuitgaven meer aan belastinginkomen zijn gekoppeld de negatieve cyclische effecten van het risico van wanbetaling op de overheidsschuld kan verminderen.

Samenvattend kunnen we stellen dat de resultaten van de analyse in de hoofdstukken 2 tot en met 5 van dit proefschrift pleiten voor een voorzichtige benadering van het fiscaal-economisch beleid, in het bijzonder wat betreft de overheidsuitgaven. De resultaten van de analyse pleiten echter niet noodzakelijk voor totale fiscale passiviteit, maar ze bevestigen in plaats daarvan de opvatting dat fiscale maatregelen niet heel effectief zijn als ze worden geïmplementeerd zonder rekening te houden met de specifieke macro-economische omgeving. Tegelijkertijd ondersteunen de resultaten het idee dat een belangrijke bijdrage van fiscaal beleid aan macro-economische stabiliteit eruit bestaat dat fiscale variabelen een duurzaam pad volgen.
Zusammenfassung | Summary in German

Fiskalpolitik und der Konjunkturzyklus
Der Einfluss von Staatsausgaben, öffentlicher Verschuldung und Staatsrisiko auf gesamtwirtschaftliche Schwankungen


1. Die empirisch orientierte makroökonomische Forschung hat bisher keinen ausreichenden Konsens zu den gesamtwirtschaftlichen Auswirkungen fiskalpolitischer Maßnahmen erreicht. Einige der Hauptgründe für den fehlenden Konsens sind erstens Schwierigkeiten bei der Unterscheidung zwischen diskretionären fiskalpolitischen Maßnahmen und automatischen Reaktionen fiskalischer Variablen wie etwa Transfers und Steuereinnahmen auf konjunkturelle Schwankungen; zweitens die Möglichkeit, dass diskretionäre Maßnahmen Wirtschaftsteilnehmern vor ihrer Implementierung bekannt sein können, zum Beispiel im Fall von angekündigten
Ausgabenprogrammen; und drittens Instabilität empirischer Ergebnisse über verschiedene Zeiträume, zum Beispiel aufgrund struktureller wirtschaftlicher Veränderungen. Die ersten beiden Probleme stellen eine wichtige Herausforderung für konventionelle empirische Methoden dar, weil sie Schätzergebnisse, die auf diesen Methoden basieren, stark verzerren können, wohingegen der Einfluss struktureller Veränderungen die Interpretation von Schätzergebnissen für überlappende Zeiträume erschwert. Ein Beispiel für eine einschlägige strukturelle Veränderung ist die vorangeschrittene weltweite wirtschaftliche Integration, die zu stärkerem Außenhandel und möglicherweise zu einer größeren Importkomponente einer fiskalpolitischen Expansion in einzelnen Ländern geführt haben könnte. Dies könnte sich wiederum auf die Effektivität einer solchen Expansion hinsichtlich deren Einfluss auf die einheimische ökonomische Aktivität auswirken.


3. Die Literatur lässt auch Spielraum für weitere Forschung zu quantitativen Kon-


Kapitel 3 zeigt, wie Daten zu Erwartungen über künftige staatliche Ausgaben verwendet werden können, um diejenigen ökonometrischen Probleme struktureller vektorautoregressiver Techniken anzugehen, die durch die Tatsache, dass diskretionäre Politikmaßnahmen oft durch Wirtschaftsteilnehmer vorhergeschen werden können, ver-
Die Zusammenfassung


Kapitel 5, welches auf gemeinsamer Arbeit mit Malte Rieth beruht, analysiert den Einfluss staatlichen Ausfallrisikos auf Konjunkturschwankungen am Beispiel eines Schwellenlands. Die Analyse basiert auf einem geschätzten quantitativen Modell für die Türkei. Das Kapitel zeigt, dass die beobachtete Dynamik der türkischen Staatsverschuldung und geldpolitisches Verhalten in Wechselwirkung mit schuldengebundenen
staatlichen Ausfallraten wichtige Aspekte vergangener Konjunkturzyklen in der Türkei erklären können. Die Ergebnisse suggerieren insbesondere, dass fiskalische Variablen Konjunkturschwankungen auch beeinflussen können (gemäß dieses Kapitels über die Effekte von schuldengebundenen staatlichen Ausfallraten auf das Konsum- und Sparverhalten von privaten Investoren), wenn die Fiskalpolitik selbst sich relativ passiv verhält. Eine mögliche Schlussfolgerung ist, dass eine stärkere Steuerbindung von Staatsausgaben die negativen zyklischen Effekte staatlichen Ausfallrisikos reduzieren könnte.

Zusammenfassend lässt sich sagen, dass die Ergebnisse, die aus der Analyse in den Kapiteln 2 bis 5 der vorliegenden Dissertation hervorgehen, einen vorsichtigen Umgang mit der Fiskalpolitik, insbesondere mit diskretionärer Staatsausgabenpolitik, über den Konjunkturzyklus befürworten. Die Ergebnisse der Analyse unterstützten jedoch nicht unbedingt totale staatliche Passivität, sondern bestätigen eher die Auffassung, dass fiskalpolitische Maßnahmen wenig effektiv sind, wenn sie ohne Berücksichtigung des spezifischen gesamtwirtschaftlichen Umfelds implementiert werden. Gleichzeitig unterstützen die Ergebnisse die Ansicht, dass ein wichtiger Beitrag staatlicher Politik zur gesamtwirtschaftlichen Stabilität ist, sicherzustellen, dass fiskalische Variablen einem nachhaltigen Pfad folgen.
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