

Household Debt and the Macroeconomy

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I Introduction

Introduction

Household debt increased substantially over the last decades. In the US economy, it rose from 96% of disposable personal income in 2000 to 128% in 2008. In some southern European countries the increase in household indebtedness was even more pronounced. In Spain, it almost doubled, from 69% in 2000 to 130% in 2008.¹

Given this significant increase in household indebtedness, recent contributions have pointed to the important role of private debt for the propagation and amplification of economic shocks and policy interventions. For example, Mian and Sufi (2011, 2012) empirically show that those US counties which experienced the largest increase in housing leverage before the financial crises, suffered from more pronounced economic slack in the postcrisis periods. They detect private debt overhang as the major reason for the slow economic recovery after the financial crisis. Within a heterogeneous agents model, Kumhof, Ranci ere, and Winant (2015) show that an increase in household indebtedness, induced by a significant rise in income inequality, makes the outburst of a financial crises more likely. Other theoretical contributions have shown the impact of fiscal policy to be larger when private indebtedness is high (Andr es, Bosc a, and Ferri, 2015; Eggertsson and Krugman, 2012; Kaplan and Violante, 2014). My thesis contributes to the still growing literature on the relation between household debt and economic activity as it tries to answer the following four questions:

1. Is there an empirical long-run relationship between income inequality and private indebtedness?
2. Is interpersonal comparison a significant determinant of short-run credit movements?
3. Do the effects of fiscal policy interventions depend on the level of private indebtedness?

¹The specific values are taken from McKinsey (2010).

4. How does private household debt evolves in response to fiscal and non-fiscal shocks?

Each of the four Chapters of my thesis focuses on one of these research questions separately. Chapters 1 and 2 study relevant determinants of households' borrowing decisions. In Chapter 1, I show that income inequality and household debt are cointegrated of order one. Thus, rising income inequality leads to a higher level of household indebtedness in the long-run. Chapter 2 provides macro-evidence for the relevance of consumption externalities between different income groups for explaining short-run credit dynamics. I estimate a business cycle model with consumer credit in which poorer households are characterized by a relative consumption motive. The keeping-up parameter is estimated to be positive and significantly different from zero indicating the important role of interpersonal comparison in understanding credit movements over the business cycle. Chapter 3 empirically studies non-linearities emerging from private debt overhang. More specifically, I show that the economic effects to fiscal consolidations crucially depend on the level of private indebtedness. When private debt is low, austerity has no significant impact on main macro aggregates. However, when private debt is high, fiscal consolidations lead to a significant and severe reduction in economic activity. In Chapter 4, I study the response of household debt to two important economic shocks: technology improvements and tax cuts. Thereby, I empirically show how household debt changes in response to both shocks, and then propose a theoretical model with financial frictions that is able to replicate the empirical responses.

The recent financial crises has been attributed to a considerable increase in income inequality by several authors (Morelli and Atkinson, 2015; Rajan, 2010). Kumhof, Rancière, and Winant (2015) study the interrelation between rising income disparity, private indebtedness, and the outburst of a financial crisis. Moreover, Iacoviello (2008) shows that the significant increase in household debt in the US economy over the last decades is closely linked to the rising income inequality observable over the same time period. Despite this growing interest and theoretical debate about the inequality-leverage nexus, the empirical research in this area is still scanty. Chapter 1 contributes

to this literature as it tests for the empirical validation of the long-run inequality-household debt relationship. Based on a panel of OECD countries, I show that income inequality and household debt are cointegrated of order one. Thus, my findings imply that rising income inequality is associated with an increase in household debt in the long-run.

Whereas Chapter 1 studies the long-run evolution of household debt, in Chapter 2, I take a closer look at the business cycle dynamics of consumer credit in the US economy. Motivated by business cycle statistics that refutes the standard consumption smoothing role of credit², I propose a business cycle model in which credit is additionally used as a source of reducing consumption disparities between different income groups. This mechanism is modeled as a consumption externality in the utility function of poorer household groups. Recent empirical studies have shown that interpersonal comparison is a significant determinant in individuals' consumption decisions (Bertrand and Morse, 2013; Carr and Jayadev, 2015; Drechsel-Grau and Schmid, 2014). I estimate deep-model parameters by matching the theoretical business cycle statistics to the empirical ones. The relative consumption parameter is estimated to be positive and significantly different from zero. This finding implies that interpersonal comparison is an important determinant of short-run credit movements. Complementary to recent microeconomic studies, this chapter provides macro-evidence on the linkage between consumption externalities and individuals' borrowing decisions.

Chapter 3 studies how private indebtedness amplifies the effects to fiscal policy interventions. Specifically, I provide empirical evidence that the consequences to fiscal consolidations crucially depend on the level of private debt overhang. Austerity measures implemented when private debt is low are hardly followed by any significant change in economic activity. In contrast, when private debt is high, fiscal consolidations lead to severe and significant reductions in private consumption and GDP. I find similar private debt-dependent effects of fiscal consolidations for other components of GDP, employment, government debt, and the default probability of the government. Two

²I show that credit is positively correlated with aggregate output and personal consumption expenditures.

central goals of fiscal consolidations are the reduction of public debt burdens and/or reducing the governments' default probability. Indeed, my results imply that consolidations implemented when private debt is high lead to a worsening of public finances and increase the probability of default. Notably, my findings are robust when controlling for two other prominent state variables: the state of the business cycle and the government debt level. Therefore, the state of the business cycle and the government debt level seem to be of minor importance for the effects of fiscal consolidations once one controls for the level of private indebtedness in the economy. I highlight two additional results detecting changes in household balance sheets as a possible transmission channel through which my findings can be rationalized. First, by differentiating between household and corporate debt, I show that most of the results are driven by household leverage. Therefore, private debt-dependent effects of fiscal policy seem to be caused by households' not firms' borrowing decisions. Second, house prices significantly decline when fiscal consolidations are implemented in high private debt states, whereas they basically do not show any effect in low private debt states. Falling house prices typically reduce the value of home equity households can use as collateral to borrow against. Chapter 3 contributes to the literature as it tests for the validity of existing theoretical models which show that private indebtedness matters for the transmission of fiscal policy (for example Andrés, Boscá, and Ferri, 2015; Eggertsson and Krugman, 2012; Kaplan and Violante, 2014). In fact, I provide extensive empirical evidence that confirms predictions of theories pointing out the impact of fiscal policy interventions to be larger in periods of private debt overhang. Moreover, my results help understanding the dismal growth performances in southern European countries, which implemented large-scale fiscal consolidation programs while confronted with high private debt levels.

In Chapter 4, I study how household debt reacts to two exogenous innovations, namely TFP shocks and tax cuts. In the empirical part of the Chapter, it is shown that both shocks induce a significant and persistent increase in household debt. Output, durable, and non-durable consumption also increase in a humped-shaped manner implying a strong comovement between household debt and aggregate economic activity.

In the second part of the Chapter, I propose a theoretical model that is able to account for the empirical responses. The model is populated by two types of household who differ in their willingness to postpone consumption into the future, creating lenders and borrowers. In contrast to the lender, the borrowing capacity of indebted households is limited to a fraction of their non-depreciated stock of durables. To match theoretical impulse responses with the empirical data, deep model parameters are estimated by the simulated method of moments approach. The results suggest that the model can successfully account for the sizes and the hump-shaped patterns of the empirical dynamics in all variables. In line with the empirical findings, the model produces persistent increases in household debt, which last for more than 20 quarters. Moreover, the debt responses almost perfectly match the empirical counterparts. The point estimates of deep model parameters are in line with findings of previous studies (see for example, Iacoviello, 2005; Mertens and Ravn, 2012). The Chapter contributes to the literature as it, first, presents empirical evidence on the conditional procyclicality of household debt, and second, shows that a representative agent model with incomplete financial markets as proposed by Iacoviello (2005) and Monacelli (2009) can successfully account for these empirical responses.

II Chapters

1 Inequality and Household Debt: a Panel Cointegration Analysis¹

ABSTRACT

This study investigates whether there exists an empirical long-run relationship between income inequality and household debt. By using panel cointegration techniques, I find that inequality and private leverage are cointegrated of order one and therefore share a common trending relation. Removing this trend by first differencing the series leads to biased inference. My results are robust to different indicators for household debt and alternative inequality measures. In the long-run, a one-percentage point increase in inequality is associated with an increase in household debt by 2% to 6%, depending on the inequality measure used.

Keywords: Income Inequality, Household Debt, Panel Cointegration.

JEL Codes: C23, D31, E25.

1.1 Introduction

Several authors have attributed the recent financial crisis of 2008/09 to a considerable rise in income inequality (e.g. Morelli and Atkinson, 2015; Rajan, 2010). Rajan (2010) argues that rising inequality in the United States pressured different governments to enact redistribution policies aimed at improving the lot of those low- and middle-income voters being left behind. The author also points out that in combination with a relaxation of underwriting standards, rising income disparity led to an increasing use of credit unsupported by greater income. The resulting credit bubble is seen as one of the foundations for the subsequent crisis (Schularick and Taylor, 2012). Following this argumentation, Kumhof, Ranci ere, and Winant (2015) study the relationship between

¹A shortened version of this chapter is published as Klein (2015).

income inequality, household debt and, the likelihood of a financial crisis within a DSGE framework. Additionally, Iacoviello (2008) develops a heterogeneous agents model which is able to capture the trend and cyclical behavior of debt and income dispersion for the US economy.

Despite this growing interest and theoretical debate about the inequality-leverage-crisis nexus, the empirical research in this area is still scanty. One exception is the study by Bordo and Meissner (2012). They explicitly test the empirical support for the hypothesis set up by Rajan (2010) and Kumhof, Ranci re, and Winant (2015) within a panel dataset covering 14 advanced economies. By estimating the effect of changes in income inequality on the change of bank loans, the authors do not find a significant relationship between inequality and bank loans growth. The results of Bordo and Meissner (2012) coincide with those of Morelli and Atkinson (2015) who fail to find a causal relation between changes in income inequality and economic crises.

My study differs from those by Bordo and Meissner (2012) and Morelli and Atkinson (2015) in three important dimensions. First, a more precise measure of household debt offered by the Bank for International Settlements (BIS) is used. The BIS debt variable measures the outstanding amount of credit to private households and therefore does not include credit to the business sector as the series used in Bordo and Meissner (2012) does. Second, in order to check for the sensitivity of the results, I consider four different inequality indicators. Three of these series, namely the top 1% income share, the inverted Pareto-Lorenz coefficient, and the Gini index measure the income distribution within one economy. The fourth one, the labor income share, includes information about the distribution of factor incomes. Most importantly, this study differs from Bordo and Meissner (2012) and Morelli and Atkinson (2015) in the underlying hypothesis tested. While Bordo and Meissner (2012) and Morelli and Atkinson (2015) investigate the relationship between *changes* in income inequality and private debt, I test for the relationship between inequality and debt in *levels*.

By using panel cointegration methods, I test whether the levels of income inequality and household debt share a common long-run relationship. Based on all these consid-

erations, my study can be seen as a more precise and general approach for testing for the existence of a long-run relationship between income inequality and household debt as hypothesized in Rajan (2010) and theoretically modeled in Kumhof, Ranci re, and Winant (2015) and Iacoviello (2008).

My results suggest that there exists a long-run relationship between income disparity and household debt. This result is robust to all four inequality measures considered. Moreover, a common trending relation is present whether the underlying cointegration test allows for cross-sectional dependence or not. Depending on the inequality indicator used, in the long-run, a one-percentage point increase in inequality leads to an increase in household debt by 2% to 6%.

The remaining chapters of the paper are organized as follows. Section 2 will review the existing literature on the connection between inequality, credit, and financial crises. Section 3 describes the panel cointegration tests used in the study. Section 4 presents the data and addresses the problem of unit root tests on bounded variables because some of the inequality measures considered in this paper have a limited value range. Cointegration test results are reported in Section 5. Estimations of the long-run relationship between inequality and household debt are presented in Section 6 and section 7 concludes.

1.2 Related Literature

Rajan (2010) proposes a linkage between inequality, credit expansion, and financial crisis in the United States in the first decade of the 21st century. Rajan argues that rising inequality led to political pressure for redistribution in the form of subsidized housing finance via institutions like Fannie Mae and Freddie Mac. The resulting lending boom created an unsustainable increase in house prices which reversed in 2007 and finally can be identified as one major reason for the crisis of 2008/09. Along these lines, Kumhof, Ranci re, and Winant (2015) model a relationship between inequality, household debt, and the probability of a crisis within a DSGE framework. Their model consists of two representative agents: an investor, who owns all of the capital, earns

only capital income, and saves and invests as well as consumes; and a worker who earns wage income, demands loans offered by the investor, and uses these income sources for consumption. A negative shock on the bargaining power of workers leads to an increase in income differences between the two agents. Due to a subsistence level of consumption included in the worker's utility function, the pronounced rise in inequality results in an increasing amount of loans demanded by workers, in order to maintain the desired level of consumption. Consequently, workers' household debt rises as well. Because the authors assume a convex relation between household debt and the probability of an economic crisis, they connect rising inequality to an increasing amount of leverage and, ultimately to a higher probability of a crisis. Ranci ere et al. (2012) extend this model to an open economy framework.

By using a heterogenous agents model, Iacoviello (2008) is able to replicate the long-run and short-run dynamics of household debt and income inequality in the United States. Based on the theoretical model set up by Krusell and Smith (1998), agents face aggregate and idiosyncratic income shocks and accumulate real and financial assets. In the model there are so-called patient agents which have a low discount rate and do not face borrowing constraints and impatient agents which discount the future more heavily and face a collateral constraint. In response to a negative idiosyncratic income shock, unconstrained agents reduce consumption by a small amount but increase their debt. Instead, constrained agents behave like hand-to-mouth consumers by reducing consumption and borrowing less. The simulated model successfully captures the observed income inequality and household debt series. Additionally, the model attributes the trend increase in debt to the pronounced rise in inequality, whereas business cycle fluctuations can account for the short-run changes in household debt.

Although models like those by Kumhof, Ranci ere, and Winant (2015) or Iacoviello (2008) explicitly make use of a connection between inequality and household debt, there is only a small literature testing for this relationship empirically. Morelli and Atkinson (2015) study the question whether economic crises were preceded by rising inequality. By using a dataset that covers 25 countries over the period from 1911 to 2010, they

do not find any relationship between changes in income inequality and banking crises. Nevertheless, they conclude that "[...] we have not investigated whether [the, note of the author] inequality level was relatively higher before identified macroeconomic shocks. Therefore, the level hypothesis cannot be ruled out at this stage." (Morelli and Atkinson, 2015, p. 49). Following this considerations, Bellettini and Delbono (2013) show that between 1982 and 2008, a large majority of banking crises have been preceded by persistently high levels of income inequality. However, Morelli and Atkinson (2015) and Bellettini and Delbono (2013) focus on the relationship between income inequality and the occurrence of a banking crisis and not on the connection between inequality and household debt which is essential in the models of Kumhof, Ranci ere, and Winant (2015) and Iacoviello (2008).

Bordo and Meissner (2012) empirically study the relationship between changes in inequality and credit growth. Based on the dataset of Schularick and Taylor (2012), they use the amount of outstanding bank loans to the private sector as an indicator for household debt. The inequality measure in their study is the share of income of the top 1%. By using panel data on 14 advanced countries for the period from 1920 to 2000, they do not find a significant relationship between inequality growth and credit changes. Instead, interest rates and GDP per capita growth are robust determinants of credit booms. However, their study suffers from several limitations in order to test for the inequality-household debt relation set up by Rajan (2010) and used in Kumhof, Ranci ere, and Winant (2015) and Iacoviello (2008). First, the theoretical frameworks by Kumhof, Ranci ere, and Winant (2015) and Iacoviello (2008) model the connection between inequality and debt of private households. By using total loans to the private sector, credit to businesses is also included in the dependent variable used by Bordo and Meissner (2012). Given an increase in bank loans to businesses, the times series of Schularick and Taylor (2012) rises, while, *ceteris paribus*, credit to the household sector stays constant. Therefore, by using time series which explicitly measure credit to the household sector, I can investigate the relationship between inequality and household debt in more detail compared to Bordo and Meissner

(2012). Second, the authors just consider one inequality measure in their study and do not check whether their results still hold when alternative inequality variables are considered. Finally and most importantly, the theoretical works by Kumhof, Ranci ere, and Winant (2015) and Iacoviello (2008) show that there exists a trending long-run relation between income inequality and household debt. By using growth rates this trend is removed and finally just short-run dynamics remain. If, however, there is a long-run relationship between inequality and household debt, using growth rates of the variables of interest may lead to biased inference on the effect of inequality on private debt (Engle and Granger, 1987; Johansen and Juselius, 1990). In addition, as pointed out by Iacoviello (2008) short-run dynamics of household debt can well be explained by business cycle fluctuations while debt and inequality are mainly connected in the long-run. Therefore, it should not be surprising that short-run changes in GDP per capita and interest rates are significant regressors in explaining loans growth as shown by Bordo and Meissner (2012). In testing for the existence of a long-run relationship between household debt and inequality both variables should be considered in levels which is possible within the cointegration approach applied in my study.

1.3 Panel Cointegration Tests

The cointegration approach which allows testing for the presence of long-run relationships among integrated variables is a popular tool in the empirical literature (Breitung and Pesaran, 2005). However, most of the tests have only low power when applied to single unit time series mainly available just after World War II (Pedroni, 2004). Due to this dilemma, it seems natural to expand the underlying sample by including additional cross-sectional data and studying cointegration relationships within a pooled time series panel. Moreover, by applying cointegration tests, I am able to consider the variables of interest measured in levels. Therefore, my approach can be seen as more precise way for studying the existence of a long-run relationship between levels of income inequality and household debt.

In the following, I present two commonly used panel cointegration tests: the Pedroni (1999, 2004) and Westerlund (2007) test.

1.3.1 Pedroni Test

Engle and Granger (1987) developed the cointegration idea for single unit time-series. The underlying test is based on an examination of the residuals of a regression performed using I(1) variables. A necessary condition for a cointegration relationship between these variables is that the residuals of the regression should be I(0). In contrast, if the residuals are I(1) then cointegration does not exist and therefore there is no long-run steady-state relation between the variables of interest. Pedroni (1999, 2004) extend the Engle-Granger residual-based approach to the panel data setting.

The Pedroni test requires to compute the residuals from the hypothesized cointegration regression. Therefore, consider the following regression

$$y_{it} = \delta_i' d_t + \beta_i x_{it} + e_{it}, \quad (1.1)$$

where $t = 1, \dots, T$ represents the time index and $i = 1, \dots, N$ stands for the cross-sectional units. d_t contains the deterministic components, which can take three different specifications. When no deterministic trend is included in equation (1.1), then $d_t = 0$, while $d_t = 1$ in the case that y_{it} is modeled with an individual constant term. Finally, for $d_t = (1, t)'$, y_{it} is modeled with an individual constant and a time trend. Note that individual specific fixed effects and deterministic trends are allowed via the parameter δ_i . Additionally, the slope coefficients β_i can vary across individuals.

Both variables of interest y_{it} and x_{it} are assumed to be I(1) for each cross-sectional unit i . Following the Engle-Granger approach, under the null hypothesis of no cointegration the error term e_{it} should also be I(1). This can be studied by first obtaining the residuals from equation (1.1), $\hat{e}_{it} = y_{it} - \hat{\delta}_i' d_t - \hat{\beta}_i x_{it}$, and then to test whether residuals are I(1) by running the auxiliary regression for every cross-section

$$\hat{e}_{it} = \rho_i \hat{e}_{i,t-1} + u_{it}$$

or

$$\hat{e}_{it} = \rho_i \hat{e}_{i,t-1} + \sum_{j=1}^{p_i} \psi_{ij} \Delta \hat{e}_{i,t-j} + v_{it},$$

where $E[u_{it}u_{js}] = 0 \forall s, t, i \neq j$ and $E[v_{it}v_{js}] = 0 \forall s, t, i \neq j$. Thus, the individual processes are assumed to be independent and identically distributed cross-sectionally, i.e. the Pedroni test does not allow for cross-sectional correlation. Pedroni (2004) suggests seven different statistics for testing the null hypothesis of no cointegration ($\rho_i = 1$). Four out of these statistics test the null hypothesis $H_0 : \rho_i = 1$ for all i , versus the alternative hypothesis $H_1^p : \rho_i = \rho < 1$ for all i , so that a common autoregressive coefficient is presumed. Pedroni calls these four tests the within-dimension or panel cointegration tests. In contrast, if the autoregressive coefficients are allowed to vary between the cross-sectional units, the null hypothesis $H_0 : \rho_i = 1$ for all i is tested versus the alternative hypothesis $H_1^g : \rho_i < 1$ for all i . Pedroni terms these remaining three tests the between-dimension or group mean panel cointegration tests. By allowing for individual specific autoregressive coefficients, the between-dimension-based statistics take one additional source of heterogeneity into account.

1.3.2 Westerlund Test

While the first generation panel cointegration tests do not allow for cross-sectional correlation, tests of the second generation explicitly consider such dependencies.

One example of a second generation panel cointegration test is the test proposed by Westerlund (2007). In contrast to the residual-based approach, Westerlund (2007) develops an error correction-based cointegration test. The null hypothesis of no cointegration is tested by inferring whether the error-correction term in a conditional panel

error-correction model is equal to zero. This test does not rely on the common factor restriction and by employing a bootstrap approach, inference is possible even under general forms of cross-sectional dependence. In addition, as simulation results in Westerlund (2007) show, the test has good small-sample properties.

The error-correction tests are based on the following data-generating process:

$$\Delta y_{it} = \delta'_i d_t + \alpha_i (y_{i,t-1} - \beta'_i x_{i,t-1}) + \sum_{j=1}^{p_i} \alpha_{ij} \Delta y_{i,t-j} + \sum_{j=-q_i}^{p_i} \gamma_{ij} \Delta x_{i,t-j} + e_{it}. \quad (1.2)$$

Again, d_t contains the deterministic components, which can take one of the three specifications already described above.

Equation (1.2) can be rewritten as

$$\Delta y_{it} = \delta'_i d_t + \alpha_i y_{i,t-1} + \lambda'_i x_{i,t-1} + \sum_{j=1}^{p_i} \alpha_{ij} \Delta y_{i,t-j} + \sum_{j=-q_i}^{p_i} \gamma_{ij} \Delta x_{i,t-j} + e_{it}, \quad (1.3)$$

where $\lambda'_i = -\alpha_i \beta'_i$. The equilibrium relationship of the system is given by $y_{i,t-1} - \beta'_i x_{i,t-1}$. Therefore, α_i captures the speed at which the system converts back to equilibrium after an exogenous shock occurred. If $\alpha_i < 0$, then error correction is present, which implies that there exists a cointegration relationship between y_{it} and x_{it} . However, if $\alpha_i = 0$, then error correction does not happen and, thus, there is no cointegration relationship. Following these considerations, Westerlund (2007) states the null hypothesis of no cointegration as $H_0 : \alpha_i = 0$ for all i . What is considered as the alternative hypothesis depends on the assumption about the homogeneity of α_i . If the α_i 's are not required to be equal for all cross-sectional units, then H_0 is tested versus the alternative hypothesis $H_1^g : \alpha_i < 0$ for at least one i . This is done by the two so called group-mean tests. A second pair of tests, so called panel tests, make the assumption that α_i is equal across all cross-sectional units i . Thus, these panel tests are designed to test H_0 versus $H_1^p : \alpha_i = \alpha < 0$ for all units i . The distinction between

panel and group-mean cointegration tests is similar for the Pedroni and Westerlund test statistics.

The group-mean tests of the Westerlund (2007) approach can be obtained by the following three steps: First equation (1.3) is estimated by least squares for each cross-sectional unit i . This leads to

$$\Delta y_{it} = \hat{\delta}_i' d_t + \hat{\alpha}_i y_{i,t-1} + \hat{\lambda}_i' x_{i,t-1} + \sum_{j=1}^{p_i} \hat{\alpha}_{ij} \Delta y_{i,t-j} + \sum_{j=-q_i}^{p_i} \hat{\gamma}_{ij} \Delta x_{i,t-j} + \hat{e}_{it}, \quad (1.4)$$

where a caret $\hat{\cdot}$ reflects estimated parameters. Note that p_i and q_i which determine the lag and lead orders, respectively, are allowed to vary across individuals. By estimating equation (1.3), \hat{e}_{it} and $\hat{\gamma}_{ij}$ are obtained. In a second step, one computes

$$\hat{u}_{it} = \sum_{j=-q_i}^{p_i} \hat{\gamma}_{ij} \Delta x_{i,t-j} + \hat{e}_{it}.$$

Based on \hat{u}_{it} and Δy_{it} , the usual Newey and West (1994) long-run variance estimators $\hat{\omega}_{ui}$ and $\hat{\omega}_{yi}$, respectively, can be constructed. These estimators are then used to obtain $\hat{\alpha}_i(1) = \hat{\omega}_{ui}/\hat{\omega}_{yi}$. In the third and last step, the group-mean tests are computed as follows:

$$G_\tau = \frac{1}{N} \sum_{i=1}^N \frac{\hat{\alpha}_i}{SE(\hat{\alpha}_i)}, \quad G_\alpha = \frac{1}{N} \sum_{i=1}^N \frac{T \hat{\alpha}_i}{\hat{\alpha}_i(1)},$$

where $SE(\hat{\alpha}_i)$ represents the usual standard error of $\hat{\alpha}_i$.

The panel tests are also computed in three separate steps. Similar to the group-mean tests, the first step is to regress Δy_{it} and $y_{i,t-1}$ on d_t , the lagged values of Δy_{it} , and the contemporaneous and lagged realizations of Δx_{it} . Following this procedure, the projection errors can be obtained

$$\Delta \tilde{y}_{it} = \Delta y_{it} - \hat{\delta}'_i d_t - \hat{\lambda}'_i x_{i,t-1} - \sum_{j=1}^{p_i} \hat{\alpha}_{ij} \Delta y_{i,t-j} - \sum_{j=-p_i}^{p_i} \hat{\gamma}_{ij} \Delta x_{i,t-j},$$

and

$$\tilde{y}_{i,t-1} = y_{i,t-1} - \hat{\delta}'_i d_t - \hat{\lambda}'_i x_{i,t-1} - \sum_{j=1}^{p_i} \hat{\alpha}_{ij} \Delta y_{i,t-j} - \sum_{j=-p_i}^{p_i} \hat{\gamma}_{ij} \Delta x_{i,t-j}.$$

By using the values for $\Delta \tilde{y}_{it}$ and $\tilde{y}_{i,t-1}$, the common error-correction parameter, α , and its standard error are estimated in a second step.

$$\hat{\alpha} = \left(\sum_{i=1}^N \sum_{t=2}^T \tilde{y}_{i,t-1}^2 \right)^{-1} \sum_{i=1}^N \sum_{t=2}^T \frac{1}{\hat{\alpha}_i(1)} \tilde{y}_{i,t-1} \Delta \tilde{y}_{it}.$$

The standard error of $\hat{\alpha}$ is given by

$$SE(\hat{\alpha}) = \left((\hat{S}_N^2)^{-1} \sum_{i=1}^N \sum_{t=2}^T \tilde{y}_{i,t-1}^2 \right)^{-1/2} \quad \text{where} \quad \hat{S}_N^2 = \frac{1}{N} \sum_{i=1}^N \hat{S}_i^2.$$

Now suppose $\hat{\sigma}_i$ denotes the estimated standard error in equation (1.4). Then \hat{S}_i is defined as $\hat{\sigma}_i / \hat{\alpha}_i(1)$.

The last step consists of computing the panel statistics as

$$P_\tau = \frac{\hat{\alpha}}{SE(\hat{\alpha})}, \quad P_\alpha = T\hat{\alpha}.$$

To account for cross-sectional dependency within the panel, a bootstrap approach based on Chang (2004) can be applied. The method consists of the following steps.

First, the least-squares regression is fitted,

$$\Delta y_{it} = \sum_{j=1}^{p_i} \hat{\alpha}_{ij} \Delta y_{i,t-j} + \sum_{j=-q_i}^{p_i} \hat{\gamma}_{ij} \Delta x_{i,t-j} + \hat{e}_{it}. \quad (1.5)$$

By using the results of equation (1.5), the vector $\hat{\omega}_t = (\hat{e}'_t, \Delta x'_t)'$ can be computed. Here \hat{e}_t and Δx_t are vectors which contain stacked observations on \hat{e}_{it} and Δx_{it} , respectively. In the next step, bootstrap samples $\omega_t^* = (e_t^{*'}, \Delta x_t^{*'})'$ are generated by sampling with replacement the centered residual vector,

$$\tilde{\omega}_t = \hat{\omega}_t - \frac{1}{T-1} \sum_{j=1}^T \hat{\omega}_j.$$

Then the bootstrap sample Δy_{it}^* is generated. This is done by first computing the bootstrap values of the composite error term, u_{it} , via

$$u_{it}^* = \sum_{j=-q_i}^{p_i} \hat{\gamma}_{ij} \Delta x_{i,t-j}^* + e_{it}^*.$$

$\hat{\gamma}_{ij}$ is obtained by the least-squares regression of equation (1.5). For a set of p_i initial values, Δy_{it}^* can then be generated recursively from u_{it}^* as follows:

$$\Delta y_{it}^* = \sum_{j=1}^{p_i} \hat{\alpha}_{ij} \Delta y_{i,t-j}^* + u_{it}^*.$$

Once again $\hat{\alpha}_{ij}$ results from the estimation of equation (1.5). In the final step y_{it}^* and x_{it}^* are generated as

$$y_{it}^* = y_{i0}^* + \sum_{j=1}^t \Delta y_{ij}^*, \quad x_{it}^* = x_{i0}^* + \sum_{j=1}^t \Delta x_{ij}^*.$$

This step requires initiation through x_{i0}^* and y_{i0}^* which are set to zero for simplicity. Following this method step by step, leads to the bootstrap sample y_{it}^* and x_{it}^* and to the bootstrapped error-correction test. Let t_1^* denote the initial bootstrap test. By repeating this procedure S times, one will obtain t_1^*, \dots, t_S^* , which represents the bootstrap distribution of the test. The null hypothesis is then rejected if the calculated sample value of the statistic is smaller than the critical value of a lower quantile (e.g. 1%) of the bootstrap distribution.

The Pedroni and Westerlund panel cointegration tests will be applied for testing for the presence of a long-run relationship between inequality and household debt.

1.4 Data and Unit Root Tests

The underlying panel of the study consists of nine industrialized countries: Australia, Canada, France, Great-Britain, Italy, Japan, Norway, Sweden, and United States. The baseline dataset covers the period from 1953 to 2008. The main data of this study are income inequality and household debt. Four different inequality series are used: the top 1% income share, the inverted Pareto-Lorenz coefficient, the labor income share, and the Gini index. The top 1 % income share and the inverted Pareto-Lorenz coefficient are taken from the World Top Incomes Database (Atkinson, Piketty, and Saez, 2011), while the Gini index data come from the University of Texas Inequality Project (Galbraith and Kum, 2005). The inverted Pareto-Lorenz coefficient measures the ratio between the average income $y^*(y)$ of individuals with income above threshold y and the threshold y (Atkinson, Piketty, and Saez, 2011). Additionally, the value of the inverted Pareto-Lorenz coefficient does not depend on the threshold y . That is, if the coefficient equals two, the average income of individuals with income above \$100,000 is \$200,000 and the average income of individuals with income above \$1 million is \$2 million. Intuitively, a higher inverted Pareto-Lorenz coefficient leads to a fatter upper tail of the income distribution. Data on the labor share of incomes are provided by the Organisation for Economic Co-operation and Development (OECD). While the top 1% income share, inverted Pareto-Lorenz coefficient, and Gini index

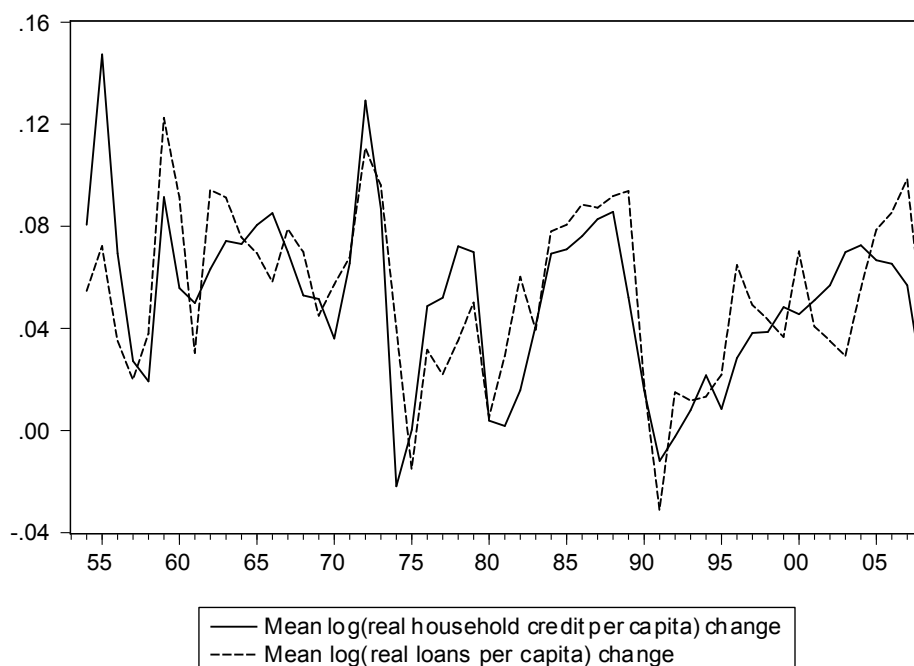
measure income distributions between persons or households, the labor income share indicates the distribution between the two factors capital and labor.

As an indicator for household debt, I use series on the outstanding amount of credit to private households and non-profit institutions serving households offered by the Bank for International Settlements (BIS). These series measure credit to the household sector and are a more precise indicator for household debt than the bank loans variable offered by Schularick and Taylor (2012) which also includes bank loans to the business sector. However, for most of the countries included in the sample, the BIS data cover only a relatively short time-span (early 1970s to 2007). Therefore, the loans series from Schularick and Taylor (2012) which is available for a longer time horizon will also be considered as a second indicator for household debt. Nevertheless, in order to accurately test for a long-run relationship between inequality and household debt, the BIS credit series will be of primary importance in the following. There are few yearly observations missing within the dataset, which are replaced by averages of the values preceding and following the missing observations.²

Figure 1.1 presents the time series of sample averages of the yearly growth rate of log of real household debt per capita based on the BIS dataset and on real loans per capita calculated from the Schularick and Taylor (2012) data. To obtain real variables, household credit as well as total bank loans are deflated by the Consumer Price index also included in the Schularick and Taylor (2012) dataset. As can be seen, at cyclical frequencies, both series move together and are strongly correlated (the overall correlation coefficient equals 0.63 and is highly significant). This observation suggests, while the BIS credit series more precisely measures credit to private households, the total bank loans variable by Schularick and Taylor (2012) follows a similar growth pattern over time.

²For more details on the data see Appendix.

Figure 1.1: Household Debt and Bank Loans



1.4.1 Unit Root Tests on Bounded Variables

In order to study whether there exists a cointegration relationship between inequality and household debt, both variables should mimic a unit root process. It seems crucial to model the data generating process for variables like the top 1% income share, labor income share, and Gini index as pure unit root processes, since ultimately these variables are bounded between the values zero and 100. It is well known that a random walk process crosses any finite bound with probability one (Jones, 1995). However, a random walk is a special case of a unit root process, namely it is linear with an additive Gaussian error (Barr and Cuthbertson, 1991). To overcome this dilemma, in the empirical literature it is preferred to think of the unit root process as a feature which describes the local behavior of the bounded series within the sample (e.g. Barr and Cuthbertson, 1991; Francis and Ramey, 2005; Guest and Swift, 2008; Herzer and Vollmer, 2012; Hurlin, 2010; Jones, 1995; Malinen, 2012; Pedroni, 2007; Young and Dove, 2013). Consequently, the unit root process is not seen as a global property but rather as a valid approximation of the underlying bounded time series within the

sample period. As pointed out by Pedroni (2007), if the determining factors of these bounded variables, such as taste, time preferences, and government policies, change over time, the series will show permanent movements that can be well described by a unit root process.

Following this line of reasoning, Pedroni (2007), Young and Dove (2013), Francis and Ramey (2005), Jones (1995), Barr and Cuthbertson (1991), and Hurlin (2010) do not reject the unit root hypothesis for several bounded variables such as investment shares, unemployment rates, bank reserve ratios, government shares of output, hours per capita, and tax rates. Herzer and Vollmer (2012), Guest and Swift (2008), and Malinen (2012) use unit root tests for studying the local behavior of different inequality measures.

Cavaliere and Xu (2014) show that conventional unit root tests tend to overreject the null hypothesis when applied to limited time series. Nevertheless, they also mention that unit root tests do not suffer from biased inference when the bounds are sufficiently far away. As all limited inequality measures considered in this study move far away from both bounds (0 and 100) and do not cross one of the bounds within the sample period, applying conventional unit root tests should not be seen as a severe problem here.

By following the aforementioned empirical literature, I approximate persistent changes in the top 1% income share, labor share of income, and Gini index as unit root processes. It seems reasonable to assume that the behavior of the bounded variables can be mimicked by a unit root data generating process (Francis and Ramey, 2005). This is done by applying two conventional panel unit root tests on the underlying inequality time series: the Fisher type ADF test as developed by Maddala and Wu (1999) and the Pesaran (2007) test. While the Maddala and Wu (1999) test belongs to the so-called first generation panel unit root tests, the test developed by Pesaran (2007) is a second generation panel unit root test (Breitung and Pesaran, 2005). The Maddala and Wu (1999) test allows for heterogeneity in the autoregressive coefficient of the Dickey-Fuller regression but ignores cross-sectional dependence in the data. In contrast, the Pesaran

Table 1.1: Panel Unit Root Tests

	Levels		First differences	
	Maddala/Wu	Pesaran	Maddala/Wu	Pesaran
Credit	16.02	1.03	40.67***	-1.61**
Loans	17.27	1.58	82.25***	-6.87***
Top 1%	6.58	-1.17	55.72***	-4.51***
Ilc	8.26	-0.63	70.47***	-1.75**
Labor share	14.47	0.36	54.06***	-1.93**
Gini	17.59	0.32	58.89***	-4.41***

All tests include individual constants and time trends. The null hypothesis is that the variable has a unit-root. *** Rejection at the 1% significant level; ** Rejection at the 5% significant level.

(2007) test assumes individual unit root processes but also allows for cross-sectional correlation in the underlying sample.

1.4.2 Unit Root Test Results

Table 1.1 presents results of the two panel unit root tests on the six variables of interest. All tests include individual constants and time trends. For all series the null hypothesis of a unit root can not be rejected when the variables are measured in levels. In contrast, when first differences are used, the Maddala and Wu test rejects the unit root hypothesis at the 1% level for all series. According to the Pesaran test statistics, first differences of the loans, top 1% income share, and Gini index reject the null hypothesis at the 1% level, while first differenced series on credit, inverted Pareto-Lorenz coefficient (Ilc), and labor share of income can be approximated as stationary processes at the 5% level. Thus, the test results do not differ significantly when cross-sectional correlation is taken into account. Cavaliere and Xu (2014) show that unit root tests tend to overreject the null hypothesis of a unit root when applied to bounded series. This finding strengthens the result of a unit root present in the limited inequality measures, as the null hypothesis can not be rejected for all relevant cases. Therefore, I conclude that both credit variables as well as the four inequality series are integrated of order one. This finding is a first prerequisite for applying cointegration tests.

1.5 Cointegration Test Results

According to the unit root test results reported in Table 1.1, stochastic trends drive the time series of both debt series and of all four inequality measures. In a next step, it will be tested if there exists a stationary linear combination between the nonstationary household debt and inequality variables, i.e. if the series are cointegrated. Two panel cointegration tests will be used. The first one is the panel cointegration test proposed by Pedroni (1999, 2004) and the second is the cointegration test developed by Westerlund (2007).

For the Pedroni test, I just report the test results applying the augmented Dickey and Fuller (ADF) principle, because, as shown in Wagner and Hlouskova (2010), those test statistics are least affected by cross-sectional correlation. In addition, these test statistics show good small sample properties (Wagner and Hlouskova, 2010). For the Westerlund test all four test statistics will be presented.

The model for testing for cointegration between inequality and household debt is:

$$\log(\text{real credit per capita})_{it} = \delta'_i d_t + \beta_i \text{inequality}_{it} + e_{it}, \quad (1.6)$$

where the level of real credit per capita is explained by the level of inequality, and $(1, -\beta_i)$ is the country-specific cointegration vector between credit and inequality. Due to heterogeneity of the data, individual constants and time trends are included in d_t . Real credit is either measured via real credit to private households from the BIS series or via real bank loans as included in the Schularick and Taylor (2012) dataset. Inequality is measured by the top 1% income share, the inverted Pareto-Lorenz coefficient, the labor income share, and the Gini index, respectively. Although some of the inequality measures are bounded, I assume that the long-run relationship between unlimited real credit per capita and possible limited inequality can well be approximated by a linear relationship as modeled in equation (1.6). This assumption is backed by the observation that all three limited inequality series move far away from their bounds

within the sample period. As found by Cavaliere (2006), standard asymptotic theory continues to provide a useful approximation when the bounds of the limited series are sufficiently far away which is the case for the inequality series considered here. Results of the panel cointegration tests based on equation (1.6) are reported in Table 1.2.

The upper part of Table 1.2 presents the results of cointegration tests based on the Pedroni (1999, 2004) ADF test statistics. While the panel ADF statistics assume a common autoregressive coefficient, group ADF statistics allow for individual specific autoregressive coefficients. Weighted panel ADF statistics refer to statistics weighted by country-specific long-run conditional variances.

19 out of the 24 test statistics reject the null hypothesis of no cointegration (at 10% level) between real credit per capita, measured as real credit to private households or real bank loans, respectively, and one of the four inequality series considered. Even at the 5% significant level, 16 out of the 24 test statistics reject the no cointegration hypothesis. When real credit to private household is used as dependent variable, the null hypothesis can be rejected at the 10% level for 10 out of the 12 test statistics. The hypothesis of no cointegration between real credit to private households per capita and the inverted Pareto-Lorenz coefficient can only be rejected for the group ADF statistics. When real bank loans are considered as a measure for private debt, the null hypothesis is rejected in nine out of the 12 cases. None of the Pedroni ADF statistics reject the no cointegration hypothesis between real bank loans per capita and the labor share at common significance levels.

The lower part of Table 1.2 reports the test results based on the Westerlund (2007) panel cointegration test which explicitly allows for cross-sectional correlation within the panel. p -values for the cointegration tests are calculated by bootstrap methods, where 800 replications are used. For each possible cointegration relationship two group mean tests (G_τ , G_α) and two panel tests (P_τ , P_α) as proposed by Westerlund (2007) are shown.

When allowing for cross-sectional dependency, the test statistics mainly support the hypothesis of cointegration between private debt and inequality. 26 out of the 32 test

Table 1.2: Panel Cointegration Test Statistics

Pedroni ADF statistics				
Dependent variable	Credit			
	Top 1%	Ilc	Labor share	Gini
Panel ADF stat	-2.14**	-0.54	-2.06**	-2.72***
weighted	-2.32***	-0.99	-2.19***	-2.71***
Group ADF stat	-2.25***	-1.89**	-1.29*	-3.91***
Dependent variable	Loans			
	Top 1%	Ilc	Labor share	Gini
Panel ADF stat	-1.97**	-2.59***	-1.01	-1.35*
weighted	-2.03**	-3.04***	-0.75	-1.28*
Group ADF stat	-1.85**	-2.52***	-0.39	-1.68**
Westerlund test statistics				
Dependent variable	Credit			
	Top 1%	Ilc	Labor share	Gini
G_τ	-2.89**	-2.86**	-2.96**	-2.25*
G_α	-12.65	-11.66	-11.04	-8.83
P_τ	-5.84*	-7.88***	-7.07**	-5.32*
P_α	-12.13*	-12.37*	-11.17*	-6.53
Dependent variable	Loans			
	Top 1%	Ilc	Labor share	Gini
G_τ	-2.99**	-3.06**	-3.29***	-2.77*
G_α	-13.80*	-13.98**	-11.32	-15.06**
P_τ	-6.86**	-7.44***	-7.15**	-6.41*
P_α	-13.50**	-14.13***	-12.85*	-14.88***

All tests include individual constants and time trends. “Weighted” refers to statistics weighted by country-specific long-run conditional variances. “ G_τ ” and “ G_α ” represent group mean test, while “ P_τ ” and “ P_α ” show panel tests. The null hypothesis is that the variables are not cointegrated. *** Rejection at the 1% significant level; ** Rejection at the 5% significant level; * Rejection at the 10% significant level.

statistics reported reject the no cointegration hypothesis at the 10% level. When real credit to private households per capita is considered as dependent variable, the null hypothesis can be rejected at the 10% level for 11 out of the 16 test statistics. If, in contrast, real bank loans per capita are considered as endogenous, 15 out of the 16 test statistics reject the no cointegration hypothesis. Cointegration between both real credit per capita measures and income disparity is present for all four inequality series considered.

When taking the findings of both cointegration tests together together, 45 out of the 56 test statistics calculated find that inequality and real credit per capita are cointegrated of order one at the 10% level. 27 out of the 32 panel test statistics and 18 out of the 24 group mean test statistics reject the no cointegration hypothesis at the 10% significant level. There are no significant differences whether real credit to private households or real total bank loans is used as measure for real credit. This finding seems surprising as it implies that including credit to the business sector in the household debt variable does not lead to different results when investigating the existence of a long-run relation between household debt and income inequality. Explaining this strong connection between total bank loans and household credit could be the subject of future research. The test results also indicate that cointegration is present for all four inequality series considered. Therefore, one can conclude that there exists a long-run relationship between inequality and household debt, i.e. that both variables have a long-run steady-state relation. This relation is present for different measures of real credit per capita and alternative inequality indicators. This finding supports the existence of a long-run relationship between inequality and household debt as modeled in Kumhof, Ranci ere, and Winant (2015) and Iacoviello (2008).

1.6 Long-run Relationship

After showing that there exists a cointegrated relationship, I want to consistently estimate the long-run effect of inequality on household debt. In doing so, the between-

Table 1.3: DOLS Estimates

	Loans	Credit
Top 1%	0.065*** (0.011)	0.064** (0.025)
Ilc	0.029** (0.012)	0.045* (0.025)
Labor share	-0.035*** (0.009)	-0.020*** (0.0008)
Gini	0.037* (0.019)	0.045* (0.024)

Standard errors are presented in parentheses. All estimations include individual constants and time trends. *** Significance at the 1% level; ** Significance at the 5% level; * Significance at the 10% level.

dimension group mean panel dynamic OLS (DOLS) estimator as proposed by Pedroni (2001) will be applied. The DOLS regression in my case is given by

$$\log(\text{real credit per capita})_{it} = \delta'_i d_t + \beta_i \text{inequality}_{it} + \sum_{j=-k_i}^{k_i} \phi_{ij} \Delta \text{inequality}_{i,t-j} + e_{it},$$

where ϕ_{ij} is a coefficient vector of lead and lag inequality differences which accounts for possible serial correlation and endogeneity of the regressor. The number of leads and lags can vary between the panel members. In the presence of cointegration the group mean panel DOLS estimator is superconsistent implying that the estimator for β converges to the true value at a faster rate than normal. The estimator is also robust to the omission of variables that do not form part of the cointegration relationship. In addition, Wagner and Hlouskova (2010) have shown that the DOLS estimator performs best in the case of cross-sectional correlation compared to several other panel cointegration estimators.

The between estimator for β is calculated as

$$\hat{\beta} = N^{-1} \sum_{i=1}^N \hat{\beta}_i,$$

where $\hat{\beta}_i$ is the conventional time-series DOLS estimator (Stock and Watson, 1993) applied to the i th unit of the panel. The associated t -statistic for the between estimator can be constructed as

$$t_{\hat{\beta}} = N^{-1/2} \sum_{i=1}^N t_{\hat{\beta}_i}.$$

The test statistics constructed from the between estimator are designed to test the null hypothesis $H_0 : \beta_i = 0$ for all i against the alternative hypothesis $H_1 : \beta_i \neq 0$. Note that under the alternative hypothesis, the values for β_i are not constrained to be the same across the panel units.

The DOLS estimates for the coefficients on the different inequality measures are reported in Table 1.3. The log of real total bank loans per capita as well as the log of real credit to private households per capita are used as dependent variable, respectively. All estimations include individual constants and time trends.

The results show that all inequality coefficients are statistically significant and have the expected signs. While the top 1% income share, the inverted Pareto-Lorenz coefficient, and the Gini index influence loans (and credit) positively, an increase in the labor share leads to a reduction in the respective dependent variable. When real loans are used as dependent variable, the absolute value of the different inequality coefficients ranges between 0.029 and 0.065, implying that, in the long-run, a one percentage point increase (one unit increase for the inverted Pareto-Lorenz coefficient) in inequality leads to a rise in real loans per capita by 2.9% to 6.5%. When real credit to private households is considered, a one percentage point (one unit) increase in inequality increases credit by 2% to 6.4%. The highest absolute coefficient values result when using the top 1% income share as regressor. The estimates are 0.065 and 0.064.

1.7 Conclusion

There is a growing interest in the relationship between income inequality, household debt, and the outburst of a financial crisis (e.g. Morelli and Atkinson, 2015; Rajan, 2010). Although in theoretical works by Kumhof, Ranci ere, and Winant (2015) and Iacoviello (2008) rising income inequality leads to an increase in household debt, there is only a small literature testing for this relationship empirically. By studying the effect of changes in income inequality on bank loans growth, Bordo and Meissner (2012) find that rises in top income shares are no significant determinant in explaining credit booms. Similar Morelli and Atkinson (2015) conclude that there is no causal relationship between rising income inequality and economic crises. However, both studies do not investigate whether there exists a relation between the *levels* of income inequality and private debt. Moreover, the models developed by Kumhof, Ranci ere, and Winant (2015) and Iacoviello (2008) explicitly use a connection between levels of income inequality and household debt. Therefore, the results by Bordo and Meissner (2012) and Morelli and Atkinson (2015) should not be seen as a rejection for the inequality-credit-crisis nexus hypothesized by Rajan (2010) and modeled by Kumhof, Ranci ere, and Winant (2015) and Iacoviello (2008).

By applying panel cointegration techniques, I have studied whether there exists an empirical long-run relation between the levels of income inequality and household debt. I have used two different measures for household debt; the private household credit series offered by the BIS and the broader Schularick and Taylor (2012) total bank loans series which also includes loans to the business sector. Additionally, four alternative inequality indicators were considered; the top 1% income share, the inverted Pareto-Lorenz coefficient, the Gini index, and the labor share of income. In testing for a cointegrated relationship between inequality and household debt, the Pedroni (1999, 2004) and Westerlund (2007) panel cointegration tests were applied. While the Pedroni test does not allow for cross-sectional correlation, a bootstrapped version of the Westerlund test makes inference under cross-sectional dependence possible.

45 out of the 56 test statistics calculated have rejected the null hypothesis of no cointegration. The results have shown no significant differences whether the household credit or total bank loans series is used as dependent variable. Additionally, the test results were robust to all four inequality measures. Therefore, it seems reasonable to conclude that there exists a long-run relationship between income inequality and leverage in developed economies which is in accordance with the theories by Iacoviello (2008), Kumhof, Ranci ere, and Winant (2015), and Rajan (2010). DOLS regressions suggest that, in the long-run, the effect of a one-percentage point increase in inequality on real loans/household credit per capita lies between 2% and 6.5%, depending on the inequality measure used.

Finally, the results by Bordo and Meissner (2012) may be considered as biased as they use first differenced variables and hence remove the long-run trend and focus on the short-term effects of changes in inequality on credit growth. Following this consideration, the finding by Bordo and Meissner (2012) is in line with Iacoviello (2008), who points out that in the short-run there is no significant relation between income inequality and household debt. At cyclical frequencies, economic activity can account for the short-run changes in household debt. In contrast, my panel cointegration results support the theoretical hypothesis of Iacoviello (2008) who finds that the long-run increase in private debt is attributed to the persistent increase in income inequality. Therefore, the cointegration approach which allows to use levels of the variables of interest seems to be more appropriate to test for the inequality-credit relation than using growth rates as done in Bordo and Meissner (2012).

1.A Appendix

The following nine countries are part of the underlying panel used in this study: Australia, Canada, France, Great-Britain, Italy, Japan, Norway, Sweden and United States.

Table A1.1: Data Definitions and Sources

Variable	Definition	Source	Missing Observations
Bank Loans	End-of-year amount of outstanding domestic currency lending by domestic banks to domestic households and non-financial corporations (excluding lending within the financial system)	Schularick and Taylor (2012)	None
Credit	Outstanding amount of credit to households and non-profit institutions serving households at the end of the year	Bank for International Settlements	None
Top 1% Income Share	Share of pre-tax household income received by the top 1 %	World Top Incomes Database	Great Britain: 1961, 1980; Italy: 1996, 1997; Norway: 1956
Inverted Pareto-Lorenz Coefficient	Ratio between the average income $y^*(y)$ of individuals with income above threshold y and the threshold y (Atkinson, Piketty, and Saez, 2011)	World Top Incomes Database	Great Britain: 1980; Italy: 1996, 1997; Norway: 1956
Labor Incomes Share	Share of national income represented by wages, salaries and benefits	OECD	None
Gini Index	Extent to which the distribution of income among households within an economy deviates from a perfectly equal distribution	Texas Inequality Project	France: 1995; Italy: 1988; United State: 2002, 2003

2 Income Redistribution, Consumer Credit, and Keeping up with the Riches

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ABSTRACT

In this study, the relation between consumer credit and real economic activity during the Great Moderation is studied in a dynamic stochastic general equilibrium model. Our model economy is populated by two different household types. Investors, who hold the economy's capital stock, own the firms and supply credit, and workers, who supply labor and demand credit to finance consumption. Furthermore, workers seek to minimize the difference between investors' and their own consumption level. We find a positive significant value for the workers' keeping up-parameter by matching business cycle statistics. Thus, our paper provides macro-evidence for the relevance of consumption externalities in explaining credit dynamics.

Keywords: Consumer Credit, Relative Consumption Motive, Business Cycles.

JEL Codes: E21, E32, E44.

2.1 Introduction

This study provides macro-evidence for the relevance of consumption externalities between different income groups. For this purpose, we propose a dynamic stochastic general equilibrium (DSGE) model with consumption externalities that is able to replicate consumer credit dynamics during the Great Moderation. By estimating deep model parameters, we show that consumption externalities are a significant determinant in explaining credit fluctuations over the business cycle. Our paper contributes to the

literatures as it integrates a well-founded mechanism into a standard DSGE model to explain short-run credit movements.

Recent empirical studies show that consumption externalities significantly affect individuals' consumption decisions. Bertrand and Morse (2013) find empirical support for so-called "trickle-down-consumption", meaning that rising income and consumption at the top of the income distribution induces households in the lower parts of the distribution to consume a larger share of their income. Focusing on the period between the early 1980s and 2008, the authors present evidence for a negative relationship between income inequality and the savings rate of middle-income households. Carr and Jayadev (2015) show that rising indebtedness of US households is directly related to high levels of income inequality. The authors conclude that relative income concerns explain a significant part of the strong increase in household leverage for the period 1999-2009. Using data from the German Socio-Economic Panel, Drechsel-Grau and Schmid (2014) demonstrate that upward looking comparison is a significant determinant of individuals' consumption decisions.

Concerning the interrelation between relative consumption concerns and private debt dynamics, there is no conclusive evidence. Bertrand and Morse (2013) provide indirect evidence that non-rich households rely on easier access to credit to finance their desired keeping up with richer co-residents. Moreover, they find that a positive relationship between the number of personal bankruptcy filings and top income levels. Georgarakos, Haliassos, and Pasini (2014) show that a higher average income increases the tendency to borrow of households with incomes below average. Contrary, Coibion et al. (2014) find that low-income households in high-inequality regions accumulate less debt than similar households in low-inequality regions. However, their findings are mainly driven by mortgages, whereas for our variable of interest, consumer credit, the authors only find mixed results. Against this background, we show within a theoretical model that relative consumption concerns are an essential driver of aggregate credit dynamics.

Understanding how unsecured consumer credit fluctuates over the business cycle is of central importance because of several reasons. First, consumer credit makes up a signif-

Table 2.1: Business Cycle Correlations (1982q1-2008q2)

	Consumer credit	
	$\sigma_{x_t}/\sigma_{D_t}$	$\rho(x_t, D_t)$
Output	0.4568	0.1523
Consumption	0.2783	0.1658
Investment	1.7524	0.0852
Hours worked	0.5080	0.3603
Real wage	0.3994	-0.3207

Note: Consumer credit has been deflated using the price index of personal consumption expenditures. All variables are logged and HP-filtered (smoothing parameter of 1600) to obtain cyclical components. Standard errors in parentheses are computed by the VARHAC-estimator with automatic lag selection by the Bayesian information criterion (see Den Haan and Levin, 1997). For data definitions and sources see Appendix.

icant part of personal consumption expenditures. For our period of interest, the Great Moderation³, credit averages 23% of aggregate personal consumption, indicating that more than one fifth of private expenditures were financed by relying on consumer credit. Second, short-run credit movements are characterized by a highly volatile behavior. As Table 2.1 reports, credit is more than twice (three times) as volatile as output (consumption). Third, and most importantly, business cycle correlations with other main aggregate variables contradict standard theory in which credit represents an instrument to smooth consumption in bad times. Table 2.1 shows positive co-movements between credit and output and consumption, respectively. Moreover, credit is positively (negatively) correlated with hours worked (real wages). The goal of this study is to show that a dynamic framework which allows for consumption externalities leads to similar credit statistics as reported in Table 2.1.

Our model economy is populated by two types of households. *Investors*, who hold the economy's entire capital stock, own firms and supply credit, and *workers*, who supply labor and demand credit to finance their desired level of consumption. Moreover, we include a mechanism through which workers value their own level of consumption relative to the investors' level of consumption, a mechanism we refer to as *keeping up*

³Following Iacoviello and Pavan (2013), and Andrés, Boscá, and Ferri (2013), among others, we date the Great Moderation as the time span between the early 1980s (here 1982q1) and the outburst of the financial crisis (2008q2).

wit the Riches. Model dynamics are driven by four stochastic innovations, namely a neutral technology, investment specific technology, price markup, and wage markup shock.

We estimate deep parameters of the four-shock model by simulated methods of moments (SMM). The parameter measuring the degree of workers' consumption externalities is estimated to be positive and significant which let us to conclude that keeping up with the riches is a central driver of credit dynamics over the business cycle. The models' implied credit moments successfully account for the (targeted) business cycle statistics as reported in Table 2.1. We also find that the estimated model replicates standard output statistics, which are not targeted in the estimation. We interpret this result as a further justification for our chosen model.

When taking a closer look at the dynamics of the estimated model version, we find that the price markup shock and the investment specific technology shock produce credit correlations which are perfectly in line with the empirical ones as reported in Table 2.1. However, this is only true when we include the consumption externality in the workers' utility function. In a counterfactual analysis we abstract from the relative consumption motive and find that the model dynamics to both shocks no more correspond to the empirical counterparts. Notably, replicating the positive correlations between credit, output, and consumption does rely on the keeping up mechanism. The neutral technology shock and the wage markup shock produce model responses that do not replicate the empirical credit correlations irrespective of the inclusion of the relative consumption motive.

The rest of the paper is organized as follows. Section 2 presents the model. In Section 3, the calibration strategy is described. Section 4 describes the models' estimation and presents its major results. In Section 5, we provide a detailed impulse response analysis of the model. It turns out that consumption externalities are of major importance for replicating credit dynamics. Finally, Section 6 concludes.

2.2 The Model Economy

This section outlines our baseline model, which consists of two types of households, a continuum of firms producing intermediate goods, a representative final good firm, and a representative labor bundler.

2.2.1 Households

The model economy is populated by a continuum of infinitely lived households, indexed on the unit interval. Following Kumhof, Ranci re, and Winant (2015), a fraction χ of households, termed as *investors* (subscript i), holds the entire stock of physical capital and owns firms, while the remaining fraction, $1 - \chi$, termed as *workers* (subscript w), makes up the entire labor force. Moreover, investors issue credit to workers. In contrast to Kumhof, Ranci re, and Winant (2015), we abstract from any default on credit. For our period of interest, the Great Moderation, delinquency rates on consumer credit in the US move around a stable mean and do not accelerate until the Great Recession. Furthermore, the respective shares of households are fixed.

Investors: Investors maximize their lifetime utility function

$$E_0 \sum_{t=0}^{\infty} \beta_i^t U_i(C_{i,t}), \quad (2.1)$$

where $\beta_i \in (0, 1)$ is the specific discount factor of investors, and $U_i(\cdot)$ is the period utility function. We assume that the level of consumption is the only argument of the investors' utility function.

Definition 1 (Investor's utility function) *We impose the following assumptions on the investors' utility function U_i .*

- (i) $\frac{\partial U_i}{\partial C_i} > 0$, $\frac{\partial^2 U_i}{(\partial C_i)^2} < 0$,
- (ii) $\lim_{C \rightarrow \infty} \frac{\partial U_i}{\partial C_i} = 0$, $\lim_{C \rightarrow 0} \frac{\partial U_i}{\partial C_i} = \infty$.

Assumption (i) states that the utility function is strictly increasing and twice differentiable in the investors' level of consumption. Assumption (ii) ensures the concavity of the utility function and that the Inada conditions hold.

The investors' budget constraint is given by

$$C_{i,t} + I_{i,t} + Q_t D_{i,t} \leq D_{i,t-1} + R_t K_{i,t-1} + \frac{\Pi_t}{\chi}, \quad (2.2)$$

where $I_{i,t}$ denotes investment, Q_t is the time t price of a credit that yields one unit of output in $t + 1$, R_t is the rental rate of capital, and Π_t/χ is the individual share of profits from ownership of firms. The law of motion for physical capital is

$$K_{i,t} = (1 - \delta)K_{i,t-1} + \zeta_t I_{i,t}, \quad (2.3)$$

where δ is the depreciation rate. ζ_t denotes a shock to the relative price of investment in terms of the consumption good. Similar to Born and Pfeifer (2014), we assume that the shock follows an AR(2)-process around its steady state value $\bar{\zeta}$,

$$\log \zeta_t = (1 - \rho_{\zeta_1} - \rho_{\zeta_2}) \log \bar{\zeta} + \rho_{\zeta_1} \log \zeta_{t-1} + \rho_{\zeta_2} \log \zeta_{t-2} + \varepsilon_{\zeta,t}, \quad (2.4)$$

where $\varepsilon_{\zeta,t} \stackrel{i.i.d.}{\sim} \mathcal{N}(0, \sigma_{\zeta}^2)$, and $|\rho_{\zeta_1} + \rho_{\zeta_2}| < 1$.

Investors maximize (2.1) subject to (2.2) and (2.3) so that the first order conditions are given by

$$\Lambda_{i,t} = U'_i(C_{i,t}), \quad (2.5)$$

$$\Lambda_{i,t} = \beta_i E_t \zeta_t \Lambda_{i,t+1} \left(R_{t+1} + \frac{1 - \delta}{\zeta_{t+1}} \right), \quad (2.6)$$

$$\Lambda_{i,t} Q_t = \beta_i E_t \Lambda_{i,t+1}. \quad (2.7)$$

Here, $U'_i(\cdot)$ denotes the first derivative of the utility function with respect to the argument in brackets, and $\Lambda_{i,t}$ denotes the Lagrange multiplier associated with (2.2). The no-Ponzi-game constraint is given by

$$\lim_{j \rightarrow \infty} E_t \frac{D_{i,t+j}}{\prod_{s=0}^j \frac{1}{Q_{t+s}}} \geq 0. \quad (2.8)$$

Workers: Each working household j maximizes the utility function

$$E_0 \sum_{t=0}^{\infty} \beta_w^t U_w(C_{w,t}, X_t, N_{w,t}(j)), \quad (2.9)$$

where $\beta_w \in (0, 1)$ is the specific discount factor of workers, $C_{w,t}$ is the workers' consumption, X_t is a consumption externality that workers take as given, $N_{w,t}$ is the individual working effort, and $U_w(\cdot)$ is the period utility function.

Definition 2 (Worker's utility function) *We impose the following assumptions on the workers' utility function U_w .*

- (i) $\frac{\partial U_w}{\partial C_w} > 0$, $\frac{\partial^2 U_w}{(\partial C_w)^2} < 0$, $\frac{\partial U_w}{\partial N_w} < 0$, $\frac{\partial^2 U_w}{(\partial N_w)^2} < 0$,
- (ii) $\frac{\partial^2 U_w}{(\partial C_w)^2} \frac{\partial^2 U_w}{(\partial N_w)^2} - \frac{\partial^2 U_w}{\partial C_w \partial N_w} > 0$,
- (iii) $\lim_{c \rightarrow \infty} \frac{\partial U_w}{\partial C_w} = 0$, $\lim_{c \rightarrow 0} \frac{\partial U_w}{\partial C_w} = \infty$,
- (iv) $\frac{\partial U_w}{\partial X} < 0$, $\frac{\partial^2 U_w}{\partial C_w \partial X} > 0$.

Assumptions (i), (ii), and (iii) refer to the standard properties of utility functions, namely that they are twice differentiable, strictly increasing in consumption, strictly decreasing in labor, strictly concave and that Inada conditions are satisfied. The first part of (iv) asserts that workers derive disutility from an increase in the consumption externality. The second part states that the marginal utility of workers' consumption is increasing in the consumption externality, implying that if this externality rises, workers wish to consume more since their marginal utility of consumption increases.⁴

⁴Including this consumption externality mechanism is backed by recent microeconomic studies, which find that upward looking comparison significantly affect individuals consumption decisions (Bertrand and Morse, 2013; Carr and Jayadev, 2015; Drechsel-Grau and Schmid, 2014).

Workers face the following budget constraint,

$$C_{w,t} + D_{w,t-1} \leq W_t(j)N_{w,t}(j) + Q_t D_{w,t} - \frac{\phi}{2}(D_{w,t} - \bar{D}_w)^2, \quad (2.10)$$

where $D_{w,t}$ denotes received credit at price Q_t , and $W_t(j)$ is the individual wage rate of household j . The last term of (2.10) represents a quadratic cost of holding a quantity of credit different from the steady state value \bar{D}_w . This assumption is needed to rule out random walk components in the equilibrium dynamics of credit.⁵ Letting $\Lambda_{w,t}$ be the workers' Lagrange multiplier on their budget constraint, the optimal choices for consumption and credit demand are determined by

$$\Lambda_{w,t} = U'_w(C_{w,t}), \quad (2.11)$$

$$\Lambda_{w,t} \left[Q_t - \phi (D_{w,t} - \bar{D}_w) \right] = \beta_w E_t \Lambda_{w,t+1}, \quad (2.12)$$

where $U'_w(\cdot)$ denotes the first derivative of the utility function with respect to the argument in brackets.

2.2.2 Final Good Firms

In this perfectly competitive sector, a representative firm produces final consumption good Y_t , combining a continuum of intermediate goods $Y_t(l), l \in [0, 1]$, using the technology

$$Y_t = \left[\int_0^1 Y_t(l)^{\frac{1}{\mu_t}} dl \right]^{\mu_t}, \quad (2.13)$$

with $\mu_t > 1$. The elasticity μ_t follows an exogenous stochastic process around its steady state value $\bar{\mu}$ given by

$$\log \mu_t = (1 - \rho_\mu) \log \bar{\mu} + \rho_\mu \log \mu_{t-1} + \varepsilon_{\mu,t}, \quad (2.14)$$

⁵Schmitt-Grohé and Uribe (2003) compare different modeling strategies that induce stationarity within small open economy models.

where $\varepsilon_{\mu,t} \stackrel{i.i.d.}{\sim} \mathcal{N}(0, \sigma_\mu^2)$, and $0 < \rho_\mu < 1$. The firm chooses intermediate inputs to maximize profits subject to (2.13), which yields the demand function for intermediate good l ,

$$Y_t(l) = Y_t \left(\frac{P_t(l)}{P_t} \right)^{\frac{\mu_t}{1-\mu_t}}, \quad (2.15)$$

and subsequently the price index of the final good,

$$P_t = \left[\int_0^1 P_t(l)^{\frac{1}{1-\mu_t}} di \right]^{1-\mu_t}. \quad (2.16)$$

2.2.3 Intermediate Good Firms

Each intermediate good is produced by a monopolistically competitive firm according to a production function given by

$$Y_t = z_t F(K_{t-1}(l), N_t(l)), \quad (2.17)$$

where we assume that F is strictly increasing, twice differentiable in both arguments, exhibits constant returns to scale, and satisfies the Inada conditions. $K_{t-1}(i)$ and $N_t(i)$ denote the quantity of capital and labor services utilized to produce intermediate good l . z_t is the technology level common across all firms and follows an exogenous stochastic process around its steady state value \bar{z} ,

$$\log z_t = (1 - \rho_z) \log \bar{z} + \rho_z \log z_{t-1} + \varepsilon_{z,t}, \quad (2.18)$$

where $\varepsilon_{z,t} \stackrel{i.i.d.}{\sim} \mathcal{N}(0, \sigma_z^2)$, and $0 < \rho_z < 1$.

Intermediate good firms maximize profits subject to the demand function (2.15) and to cost minimization. We assume identical firms and that prices are perfectly flexible so

that marginal costs are equal to $1/\mu_t$. Thus, the aggregate wage rate can be expressed as a function of the marginal product of labor, MPL_t , and μ_t ,

$$W_t = \frac{MPL_t}{\mu_t}. \quad (2.19)$$

Also, the aggregate rental rate on physical capital equals

$$R_t = \frac{MPK_t}{\mu_t}, \quad (2.20)$$

where MPK_t measures the marginal product of capital. In the context of monopolistic competition, μ_t is also known as the price markup.

Since workers make up the entire labor force, a positive shock to the price markup shifts income from workers to investors. Thus, we refer to (2.14) as a *redistribution shock*.⁶ Following Chari, Kehoe, and McGrattan (2007), among others, μ_t can also be interpreted as the labor wedge on the firm side, as it drives a wedge between the wage rate and the marginal product of labor.

2.2.4 Employment Agencies

As in Erceg, Henderson, and Levin (2000), we assume that each working household j is a monopolistic supplier of a differentiated labor service, $N_{w,t}(j)$. A representative labor bundler, termed as *employment agency*, combines the intermediate labor services into a homogenous labor input, $N_{w,t}$, using the technology

$$N_{w,t} = \left[\int_0^1 N_{w,t}(j)^{\frac{1}{\nu_t}} dj \right]^{\nu_t}, \quad (2.21)$$

with $\nu_t > 1$. The elasticity ν_t follows an exogenous stochastic process around its steady state value $\bar{\nu}$,

$$\log \nu_t = (1 - \rho_\nu) \log \bar{\nu} + \rho_\nu \log \nu_{t-1} + \varepsilon_{\nu,t}, \quad (2.22)$$

⁶Throughout the paper, we use the two terms *redistribution shock* and *price markup shock* interchangeably.

where $\varepsilon_{\nu,t} \stackrel{i.i.d.}{\sim} \mathcal{N}(0, \sigma_\nu^2)$, and $0 < \rho_\nu < 1$. The labor bundler operates in a perfectly competitive market such that profit maximization given (2.21) leads to the labor demand function

$$N_{w,t}(j) = N_{w,t} \left(\frac{W_t(j)}{W_t} \right)^{\frac{\nu_t}{1-\nu_t}}, \quad (2.23)$$

where W_t is the aggregate wage rate. By substituting (2.23) into (2.21), we obtain the following expression for the latter,

$$W_t = \left[\int_0^1 W_t(j)^{\frac{1}{1-\nu_t}} dj \right]^{1-\nu_t}. \quad (2.24)$$

We assume symmetric working households and, as for the final good price, that wages are perfectly flexible. Thus, the wage rate is defined as a function of the marginal rate of substitution, MRS_t , and the wage markup, ν_t ,

$$W_t = \nu_t MRS_t. \quad (2.25)$$

In close analogy to the price markup, ν_t can be interpreted as the labor wedge on the household side. In a perfectly competitive economy, μ_t and ν_t would be one such that wages equal the marginal product of labor on the one hand, and the marginal rate of substitution on the other.

2.2.5 Aggregation and Market Clearing

Aggregates are defined as the weighted average of the respective variables for each household type. Hence, we get

$$C_t = \chi C_{i,t} + (1 - \chi) C_{w,t}, \quad (2.26)$$

$$K_t = \chi K_{i,t}, \quad (2.27)$$

$$I_t = \chi I_{i,t}, \quad (2.28)$$

$$N_t = (1 - \chi) N_{w,t}. \quad (2.29)$$

Credit market clearing requires that

$$(1 - \chi)D_{w,t} = \chi D_{i,t}, \quad (2.30)$$

while the aggregate resource constraint is given by

$$Y_t = C_t + I_t + (1 - \chi)\frac{\phi}{2}(D_{w,t} - \bar{D}_w)^2. \quad (2.31)$$

A competitive rational expectations equilibrium is a stochastic set of sequences $\{C_t, C_{i,t}, C_{w,t}, D_{i,t}, D_{w,t}, I_t, I_{i,t}, K_t, K_{i,t}, \Lambda_{i,t}, \Lambda_{w,t}, N_t, N_{w,t}, \Pi_t, Q_t, R_t, W_t, Y_t\}_{t=0}^{\infty}$ satisfying the households' and firms' first-order conditions, as well as aggregation identities, market clearing conditions, and no-Ponzi-game constraints, given the exogenous realizations of $\{\zeta_t, \mu_t, z_t, \nu_t\}_{t=0}^{\infty}$. The model is solved by a log-linear approximation around its deterministic steady state.

2.3 Calibration

2.3.1 Functional Forms

Investors' preferences are given by

$$U_i(C_i, D_i) = \frac{C_i^{1-\sigma}}{1-\sigma}, \quad (2.32)$$

where σ is the inverse of the elasticity of intertemporal substitution. The workers' period utility is given by

$$U_w(C_w, C_i, N_w) = \frac{C_w^{1-\sigma}}{1-\sigma} X^{b\sigma} - \frac{\gamma N_w^{1+\eta}}{1+\eta}, \quad (2.33)$$

where b indicates the strength of the consumption externality, γ is a scaling parameter, and η is the inverse Frisch elasticity of labor supply. This specification implies that $MRS_t = \gamma N_{w,t}^\eta / \Lambda_{w,t}$.

X is assumed to be the contemporaneous consumption level of investors relative to the contemporaneous consumption level of workers, $X_t = C_{i,t}/C_{w,t}$. Adapting the specification of Dupor and Liu (2003), we model b as a “jealousy” parameter (i.e. $b \geq 0$), implying that an increase in the investors’ consumption level leads to a decrease in the workers’ utility level.⁷

Intermediate good firms produce according to the Cobb-Douglas production function

$$Y_t(i) = z_t K_{t-1}(i)^\alpha N_t(i)^{1-\alpha}, \quad (2.34)$$

where $\alpha \in [0, 1]$ measures the capital income share. This specification implies that $MPL_t = (1 - \alpha)Y_t/N_t$ and $MPK_t = \alpha Y_t/K_{t-1}$.

2.3.2 Parameterization

Table 2.2 shows the parameter values of the models’ baseline calibration, where an upper bar denotes the steady state value of the respective variable. The simulated data of the model are at a quarterly frequency. The depreciation rate of capital, δ , equals 2.5%, which corresponds to an annual depreciation rate on capital equal to 10 percent. The discount factor of both agents is set to 0.995, which, combined with δ , implies a real interest rate on capital of 3%.

Following the empirical study by Bertrand and Morse (2013), the share of investors (rich households) in the overall population, χ , is set to 20%. We normalize the steady state level of labor to 0.33 and set the inverse Frisch elasticity, η , to 1, which is in the range of values suggested by Hall (2009). The capital share parameter, α , equals 0.27. The steady state levels \bar{z} and $\bar{\zeta}$ are normalized to 1.

In equilibrium, marginal cost equals the inverse of the price markup. The steady state values for the rental rate of capital and marginal costs are used to calculate the steady state wage rate, which leads, subsequently, to the steady state level of capital. After obtaining the levels of both input factors, we are able to calculate the steady

⁷Similar specifications of relative consumption motives are used by Alvarez-Cuadrado and El-Attar (2012) and Al-Hussami and Remesal (2012) who study the effect of rising inequality on individual saving rates and current account imbalances, respectively.

Table 2.2: Model Calibration

	Parameter	Value
Preferences		
Discount factors	$\beta_i = \beta_w$	0.995
Inverse Frisch elasticity	η	1
Fraction of investors	χ	0.20
Technology		
Capital share	α	0.27
Capital depreciation rate	δ	0.025
Steady state		
Price markup	$\bar{\nu}$	1.1
Wage markup	$\bar{\mu}$	1.1
Labor	\bar{N}	0.33
Credit-to-labor income	$\bar{D}_w / (\bar{W} \bar{N}_w)$	0.27
Neutral technology	\bar{z}	1
Investment specific Technology	$\bar{\zeta}$	1

state output level. The resulting steady state investment-to-GDP ratio equals 17%, which is in line with US data for our sample period.

Because the price markup is larger than unity, profits are positive in equilibrium. We set $\bar{\mu}$ and $\bar{\nu}$ to 1.1 so that steady state markups in the product and labor market are 10%, which is in the range of values typically used in the literature. By assuming a steady state consumer credit-to-labor income ratio for workers, $\bar{D}_w / (\bar{W} \bar{N}_w)$, of 27%, which is the average for the Great Moderation, and using the two budget constraints, we determine the consumption levels of both agents. The investors' consumption level is then used to obtain the respective shadow price of consumption $\bar{\lambda}_i$.

The workers' shadow price of consumption depends on the parameter b , which measures the strength of the relative consumption motive in their utility function. In what follows, we estimate b and other deep model parameters by SMM. Finally, after obtaining b , the parameter γ is calibrated via the steady state labor supply condition.

2.4 Model Estimation

We estimate the characteristics of the technology shock and the redistribution shock by ordinary least squares (OLS). Due to data limitations, the remaining shock parameters are estimated with a Simulated Method of Moments (SMM) approach. Moreover, the most crucial parameters for our model, namely σ and b , which determine the impact of the relative consumption motive, are also included in this estimation procedure.

2.4.1 OLS Estimation

As observation period, we select the Great Moderation, ranging from 1982q1 to 2008q2. With the exception of the technology shock series, all data series mentioned in the following are obtained from the FRED database.

Data on the technology shock are taken from Fernald (2012). The variable is detrended before estimation by a one-sided HP-filter with a smoothing value of 1600 as suggested by Born and Pfeifer (2014). The estimated AR-coefficient and standard deviation are 0.837 and 0.008 respectively. These estimates are similar to the findings of Bullard and Singh (2012).

For constructing a time series of the redistribution shock, we follow Galí, Gertler, and López-Salido (2007) and use the following equation,

$$\mu_t = MPL_t - w_t, \tag{2.35}$$

where the marginal product of labor, MPL_t , equals $\log[(1 - \alpha)y_t/n_t]$. y_t/n_t is measured as the real output per hour worked of all persons in the nonfarm business sector, and w_t is the log of real compensation per hour in this sector. Again, all series are detrended by the one-sided HP-filter. The estimates of the AR-coefficient and the standard deviation are 0.777 and 0.006 respectively, and thus, similar to those of Galí, Gertler, and López-Salido (2007) and Karabarbounis (2014). The upper part of Table 2.3 summarizes the parameter values estimated by OLS.

2.4.2 SMM Estimation

According to (2.25), the wage markup, ν_t , is defined as the product of the real wage rate, W_t , and the marginal rate of substitution, MRS_t . Given the specific utility function of workers,

$$MRS_t = \frac{\gamma N_{w,t}^\eta}{\Lambda_{w,t}}, \quad \text{where} \quad \Lambda_{w,t} = C_{w,t}^{-\sigma} X^{b\sigma}, \quad (2.36)$$

to calculate a wage markup series, we would need data on $C_{i,t}$ and $C_{w,t}$, and an appropriate value for b , the parameter measuring the strength of the relative consumption motive. However, since there is no such data available to the best of our knowledge and there is little guidance in the literature about values for b , we use the SMM estimator, originally proposed by McFadden (1989) and Lee and Ingram (1991), to overcome the data problem. The objective of SMM is to find a parameter vector that minimizes the weighted distance between simulated model moments and their empirical counterparts.

Let $\widehat{\Omega}$ be a $k \times 1$ vector of empirical moments computed from the data and let $\Omega(\theta)$ be the $k \times 1$ vector of simulated moments computed from artificial data. The corresponding time series are generated from simulating the model given a draw of random shocks and the $p \times 1$ vector $\theta \in \Theta$, with $\Theta \subseteq \mathbb{R}^p$. The length of the simulated series is τT , where T is the number of observations in the real data set and $\tau \geq 1$ is an integer. Then, the SMM estimator is given by

$$\tilde{\theta}_{SMM} = \arg \min_{\theta \in \Theta} \left[\widehat{\Omega} - \Omega(\theta) \right]' W^{-1} \left[\widehat{\Omega} - \Omega(\theta) \right], \quad (2.37)$$

where W is a $k \times k$ positive-definite weighting matrix.

Specifically, $\widehat{\Omega}$ contains the consumer credit moments as shown in Table 2.1. $\tilde{\theta}_{SMM}$ contains the estimates for b , σ , ϕ , $\rho_{\zeta,1}$, $\rho_{\zeta,2}$, σ_ζ , ρ_ν , and σ_ν . For the weighting matrix, we follow Ruge-Murcia (2013) and choose a matrix with diagonal elements equal to the optimal weighting matrix while all off-diagonal elements are equal to zero.⁸ Hence, we

⁸Ruge-Murcia (2013) shows that this choice produces consistent parameter estimates, while standard errors are just slightly higher than those generated with the optimal weighting matrix. The

only put weight on moments that are observed in the data and force the estimation to consider only economically meaningful moments (see Cochrane, 2005, chap. 11). Additionally, we follow Born and Pfeifer (2014) and incorporate prior information about the parameters to estimate. In particular, we choose prior means $\bar{\theta}$ for each parameter in θ and expand $[\hat{\Omega} - \Omega(\theta)]$ by $(\tilde{\theta}_{SMM} - \theta)$, the deviation of the estimated parameter from the respective prior mean. We expand W by attaching small penalty terms to the diagonal, which raise the objective function when deviating from the prior mean. We choose this procedure to rule out local minima in implausible regions of the state space which is often the case when estimating DSGE models.⁹ We choose a prior mean of 0 to be agnostic about the strength of the relative consumption motive b .

To rule out dependence on one particular draw of shocks, we draw several sets of shocks and choose the parameter set that minimizes the average objective function. We use the following algorithm to estimate θ .

Algorithm 2.1 (Construction of objective function to be minimized) *We start with a guess for $\tilde{\theta}_{SMM}$. Then:*

1. *Draw 50 sets of shocks, each consisting of $(\tau T + 1500) \times 4$ values.*
2. *For each set of shocks: solve the model, simulate time series, discard the first 1500 periods, compute moments, compute objective function.*
3. *Take average over all 50 objective function values and minimize this.*

For the minimization, we use the nonlinear optimization routine proposed by Byrd, Nocedal, and Waltz (2006). All parameters are set as in the baseline calibration (see Table 2.2), except for those of $\tilde{\theta}_{SMM}$. We set τ to 10, implying that the artificial time series are ten times larger than the original sample size. Ruge-Murcia (2013) shows that this is a useful choice for handling the trade-off between accuracy and computational cost.

optimal weighting matrix is given by the inverse of the variance-covariance matrix associated with the sample moments.

⁹Also known as the “dilemma of absurd parameter estimates”, see An and Schorfheide (2007).

Following Ruge-Murcia (2013), we compute the standard errors of $\tilde{\theta}_{SMM}$ from an estimate of its asymptotic covariance matrix as

$$(1 + 1/\tau)(J'WJ)^{-1}J'WJSJ(J'WJ)^{-1}, \quad (2.38)$$

where J is the Jacobian matrix and S is the full variance-covariance matrix of the empirical moments.

Table 2.3: Parameter Values for Model Simulation

	Parameter	Value	SD
OLS estimation			
AR(1)-coefficient technology shock	ρ_z	0.8368	(0.0554)
Standard deviation technology shock	σ_z	0.0084	(0.0031)
AR(1)-coefficient redistribution shock	ρ_μ	0.7769	(0.0629)
Standard deviation redistribution shock	σ_μ	0.0063	(0.0024)
	Parameter	Value	SD
SMM estimation			
Relative consumption motive	b	6.5231	(0.2776)
Inverse substitution elasticity	σ	1.5368	(0.0921)
Debt adjustment cost parameter	ϕ	1.3178	(0.0104)
AR-coefficient investment-specific technology shock	$\rho_{\zeta,1}$	0.9045	(0.3230)
AR-coefficient investment-specific technology shock	$\rho_{\zeta,2}$	-0.0192	(0.3410)
Standard deviation investment shock	σ_ζ	0.0141	(0.0017)
AR-coefficient wage markup shock	ρ_ν	0.4671	(0.1203)
Standard deviation wage markup shock	σ_ν	0.0272	(0.0040)

The results of the SMM estimation are shown in the lower part of Table 2.3. For b , we obtain a value of 6.523 which is estimated to be significantly different from zero, indicating a strong presence of the relative consumption motive. To get a better interpretation of this value, we directly relate this estimate to the findings of Bertrand and Morse (2013). In doing so, we simulate an exogenous increase in the investors' income stream and compare the implied investors' consumption response to the respective workers' consumption change. A 1% increase in consumption by investors is associated with 0.8% higher consumption expenditures by working households. This elasticity is in the upper range of estimates provided by Bertrand and Morse (2013),

Table 2.4: Data and Simulated Model Correlations

	$\sigma_{x_t}/\sigma_{D_t}$		$\rho(x_t, D_t)$		$\sigma_{x_t}/\sigma_{Y_t}$		$\rho(x_t, Y_t)$	
	Data	Model	Data	Model	Data	Model	Data	Model
Output	0.4568	0.3615	0.1523	0.1272	-	-	-	-
Consumption	0.2783	0.3034	0.1658	0.1970	0.6092	0.8395	0.8020	0.7529
Investment	1.7524	1.1733	0.0852	-0.0063	3.8359	3.2472	0.9061	0.7489
Hours worked	0.5080	0.5479	0.3603	0.3326	1.1120	1.5167	0.8144	0.7342
Real wage	0.3994	0.4541	-0.3207	-0.4386	0.8743	1.2576	0.0023	-0.2970

Note: Consumer credit has been deflated using the price index of personal consumption expenditures. All variables are logged and HP-filtered (smoothing parameter of 1600) to obtain cyclical components. For data definitions and sources see Appendix.

which implies that our estimated model is able to replicate microeconomic evidence on the strength of the keeping up motive.

The values for σ , ϕ , $\rho_{\zeta,1}$ and $\rho_{\zeta,2}$ are in the range of values typically found in other studies. The AR-coefficient for the wage markup shock displays a relatively low degree of persistence with a value of 0.467. Moreover, the standard deviations of both shocks, σ_{ζ} and σ_{ν} , are in line with values found by related studies.

Columns 2-5 of table 2.4 report the credit moments obtained from the data and from the model simulation. All these model moments are close to those in the data with only minor discrepancies. In line with the empirical observation, output, consumption, hours worked, and the real wage are less volatile than consumer credit, whereas investment show a higher volatility. As in the data, the models' responses imply positive correlations between consumption, output, hours worked, and credit whereas the real wage and consumer credit are negatively correlated. Additionally, investment show no clear correlation with the credit cycle. The rather negligible differences suggest that our calibration/estimation exercise provides a set of reasonable parameter values and, furthermore, supports the inclusion of the *keeping up with the Riches* mechanism.

Columns 6-9 reveal the correlations between output and the remaining four measures. Note that these coefficients were not included in the moment-matching approach. Thus, we interpret these results as the model's ability to replicate important conventional business cycle relations.

Simulating the model leads to a strong procyclical behavior of investment and hours worked with coefficients close to the empirical moments. The model is also able to produce a strong positive co-movement between output and consumption as observed in the data. While the wage rate is completely acyclical in the data, the two series are negatively correlated in the model simulation. However, the differences between the two sets of moments are only small-sized so that we interpret the results of this quantitative exercise as a validation of our proposed model and the underlying calibration/estimation strategy.

2.5 Impulse Responses

In Section 2.1, we have shown that for the period of the Great Moderation consumer credit exhibits

- a) positive correlations with consumption and output,
- b) a positive correlation with hours worked,
- c) a negative correlation with real wages.

In this section, we use our proposed model to study the effects of a neutral technology shock, a wage markup shock, a price markup shock, and an investment-specific technology shock and assess their ability to reproduce these empirical relationships. All parameter values are set according to Table 2.2 and Table 2.3.

As mentioned above, the two markup shocks are closely related to the labor wedge, which is responsible for substantial unexplained cyclical fluctuations.¹⁰ By including these shocks, we stress the importance of the labor wedge not only for labor market outcomes but also for the behavior of consumer credit over the past three decades.

We present model responses for two different values for b , the estimated value \hat{b} and $b = 0$, while holding all remaining parameter values constant, to highlight the impact of the workers' relative consumption motive.

¹⁰See Hall (1997), Galí, Gertler, and López-Salido (2007), Chari, Kehoe, and McGrattan (2007), Shimer (2009), Karabarbounis (2014).

2.5.1 Neutral Technology Shock

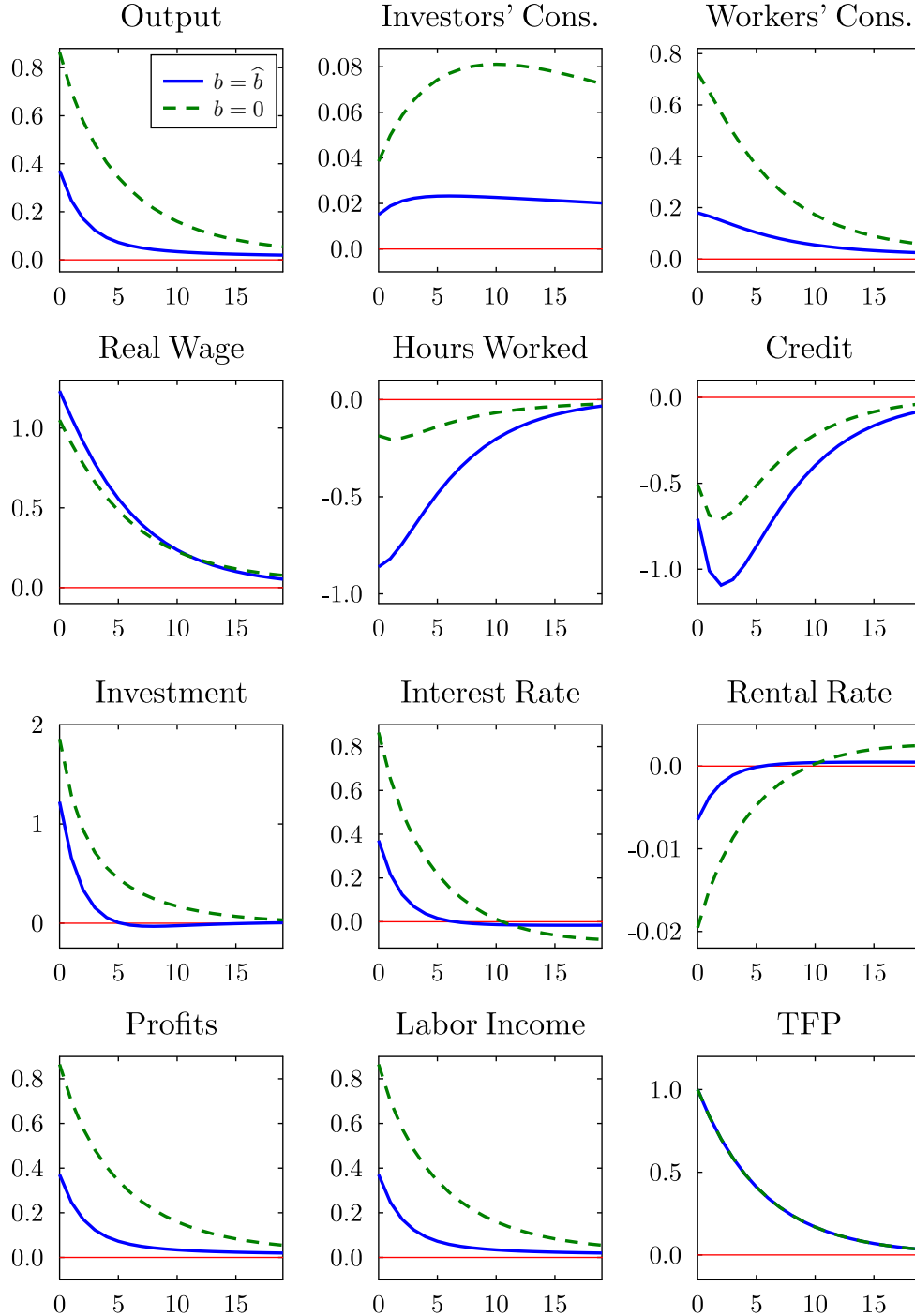
Figure 2.1 shows the effects of a positive neutral technology shock to the model economy. We first discuss the results for $b = \hat{b}$ (solid lines). An increase in z_t causes output to go up immediately. As a result of the rise in productivity, the marginal products of labor and capital increase, leading to a higher wage rate and interest rate on capital. Both agents increase their respective consumption levels, although the rise is more pronounced for working households. As $\hat{b} > 0$, workers minimize consumption differences by reducing hours worked and credit demand substantially. However, workers' total labor income increases as the rise in the wage rate is more pronounced than the fall in hours worked. Real profits increase by a similar magnitude as output.

If we abstract from the relative consumption motive, $b = 0$, the results (dash-dotted lines) are quantitatively different but do not change qualitatively. For $b = 0$, profits and, therefore, investors' income and consumption increase by a larger amount compared to the case of $b = \hat{b}$. Since workers now do not seek to minimize the difference between both consumption levels, they also consume more than in the case of $b = \hat{b}$. Consequently, workers reduce labor supply by a smaller amount and, in addition, reduce credit demand less strongly. As a result, workers' labor income increases stronger. Also, investment rises by a larger amount. Consequently, the rise in output to a neutral technology shock is amplified when $b = 0$.

To summarize, irrespective of the inclusion of the relative consumption motive, the neutral technology shock is able to reproduce the negative correlation between consumer credit and wages and leads to a positive co-movement between credit and labor. Nevertheless, the model generates a negative relation between consumer credit, consumption, and output which is in contrast to the data.¹¹

¹¹Here and in the following, the mentioned correlations does solely correspond to the impact responses of the respective variables.

Figure 2.1: Impulse Responses to Neutral Technology Shock



Note: Responses are measured in percentage deviations from steady state. Horizontal axes measure time in quarters.

2.5.2 Wage Markup Shock

In Figure 2.2, the effects of a positive wage markup shock are presented. For $b = \hat{b}$ (solid lines), the shock leads to a boost in the wage rate, whereas the marginal product of labor remains unchanged. Due to cost minimization, the demand for labor falls. This reduction in labor demand is so strong that, although wages rise, workers' labor income declines. Consumption smoothing combined with the relative consumption motive forces workers to demand a higher amount of credit. As the interest rate on capital decreases, investment declines as well. Combined with the falling labor demand, output decreases immediately, which leads to lower profits received by investors. Consequently, investors reduce their consumption level by a small amount. Seeking to minimize consumption differences, working households decrease their consumption expenditures as well.

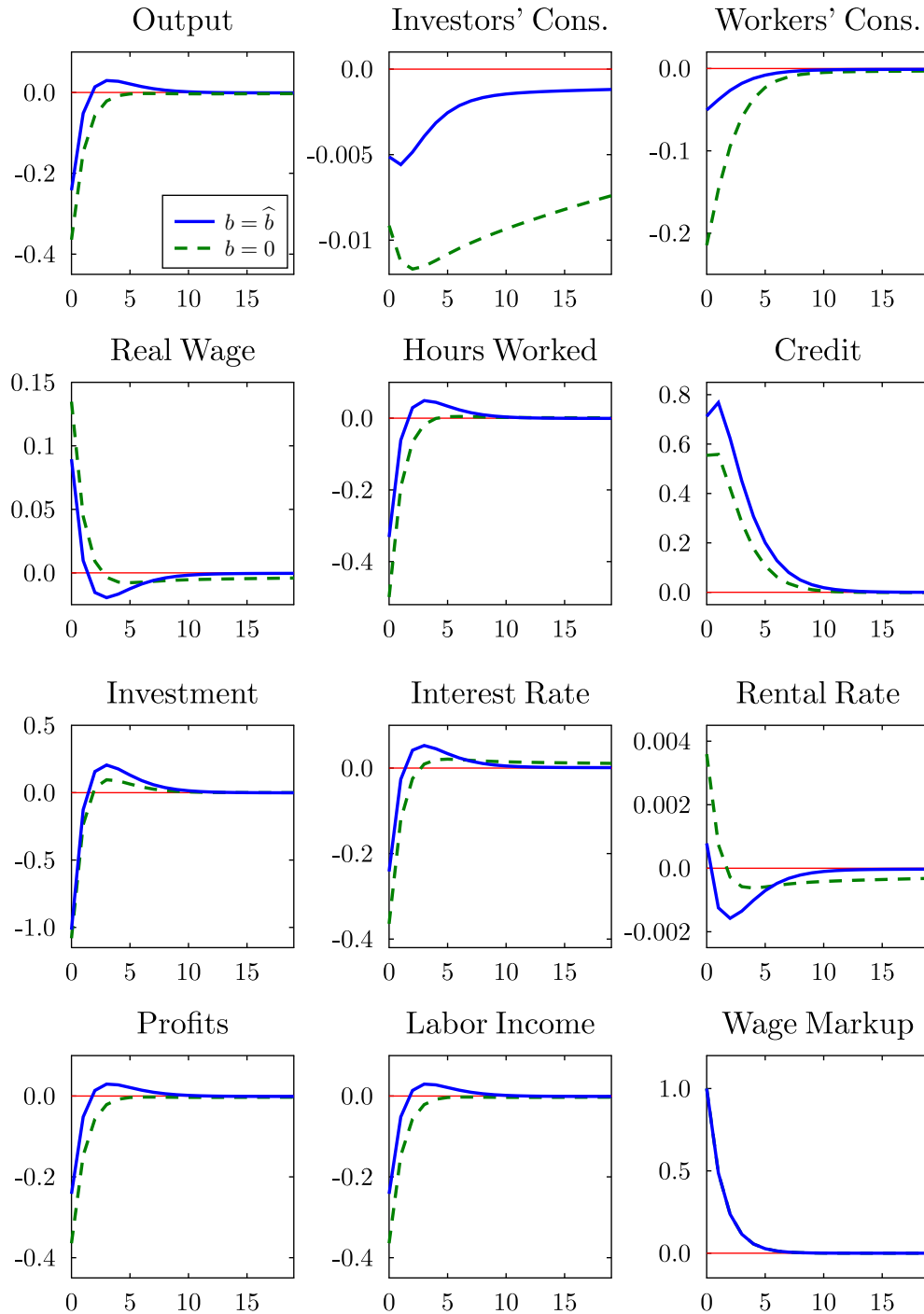
When $b = 0$, the results show only quantitative differences. The downturns in hours worked, labor income and output are more severe when we abstract from consumption externalities. Similarly, profits fall by a larger amount such that investors' consumption level decreases stronger. In addition, the reduction in workers' consumption is larger when the relative consumption motive is not present.

In contrast to the data, the model generates a positive correlation between consumer credit and wages as a response to a wage markup shock and negative co-movements between consumer credit, consumption and labor. These results do not depend on the presence of the relative consumption motive in the workers' utility function.

2.5.3 Price Markup Shock

Figure 2.3 presents the model responses to the price markup shock. The shock leads to a falling wage rate, while not affecting the marginal product of labor. A similar effect can be observed for the rental rate of capital. Due to lower marginal cost, profits rise such that investors obtain a higher income and increase their consumption level. If the relative consumption motive is present ($b = \hat{b}$, solid lines), which induces workers to minimize consumption differences, working households respond to the rise in investors

Figure 2.2: Impulse Responses to Wage Markup Shock



Note: Responses are measured in percentage deviations from steady state. Horizontal axes measure time in quarters.

consumption by, first, increasing their labor supply. As a result, the absolute drop in labor income is smaller than the wage reduction. Second, workers enhance their demand for credit to finance their desired level of consumption. Both effects induce an increase in workers' consumption expenditures. As investment and hours worked rise, aggregate output also goes up when the price markup shock hits the economy.

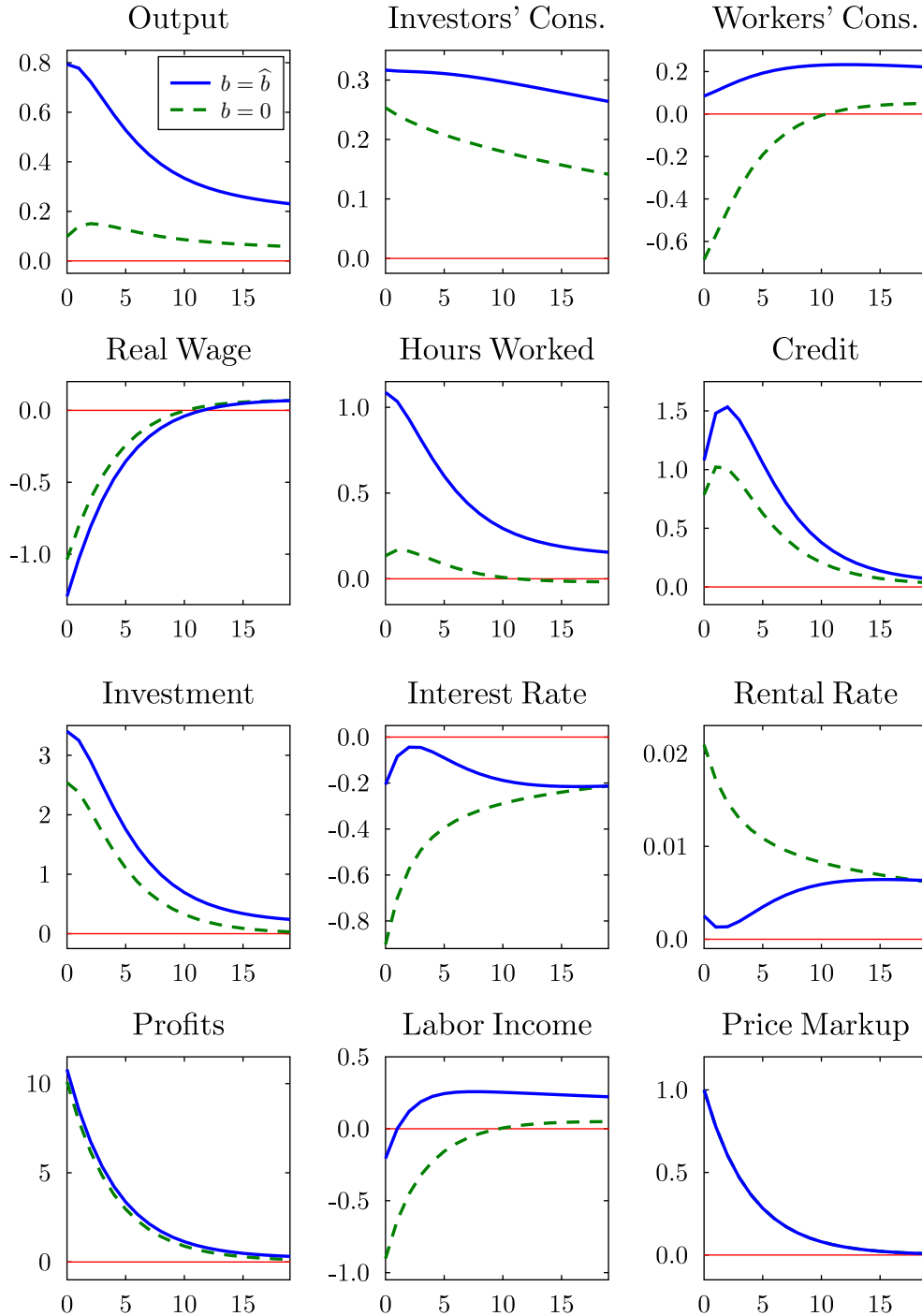
The situation changes if we abstract from the relative consumption motive ($b = 0$), so that the workers' choice of consumption only depends on consumption smoothing. In this case, workers increase their labor supply by a smaller amount. As a result, the drop in labor income is more pronounced. Consequently, also output goes up by a smaller amount. Although investors still consume more than in steady state, workers reduce their consumption expenditures when abstracting from the relative consumption motive.

If $b = 0$, a price markup shock produces a negative correlation between consumer credit and wages and positive correlations between credit, labor and output, in line with the empirical counterparts. However, the implied model correlation between credit and consumption is negative which stands in contrast to the data. Nevertheless, if the consumption externality is present, the model produces a positive relation between credit and consumption. Thus, if $b = \hat{b}$ the models' responses to a price markup shock are perfectly in line with the data.

2.5.4 Investment Specific Technology Shock

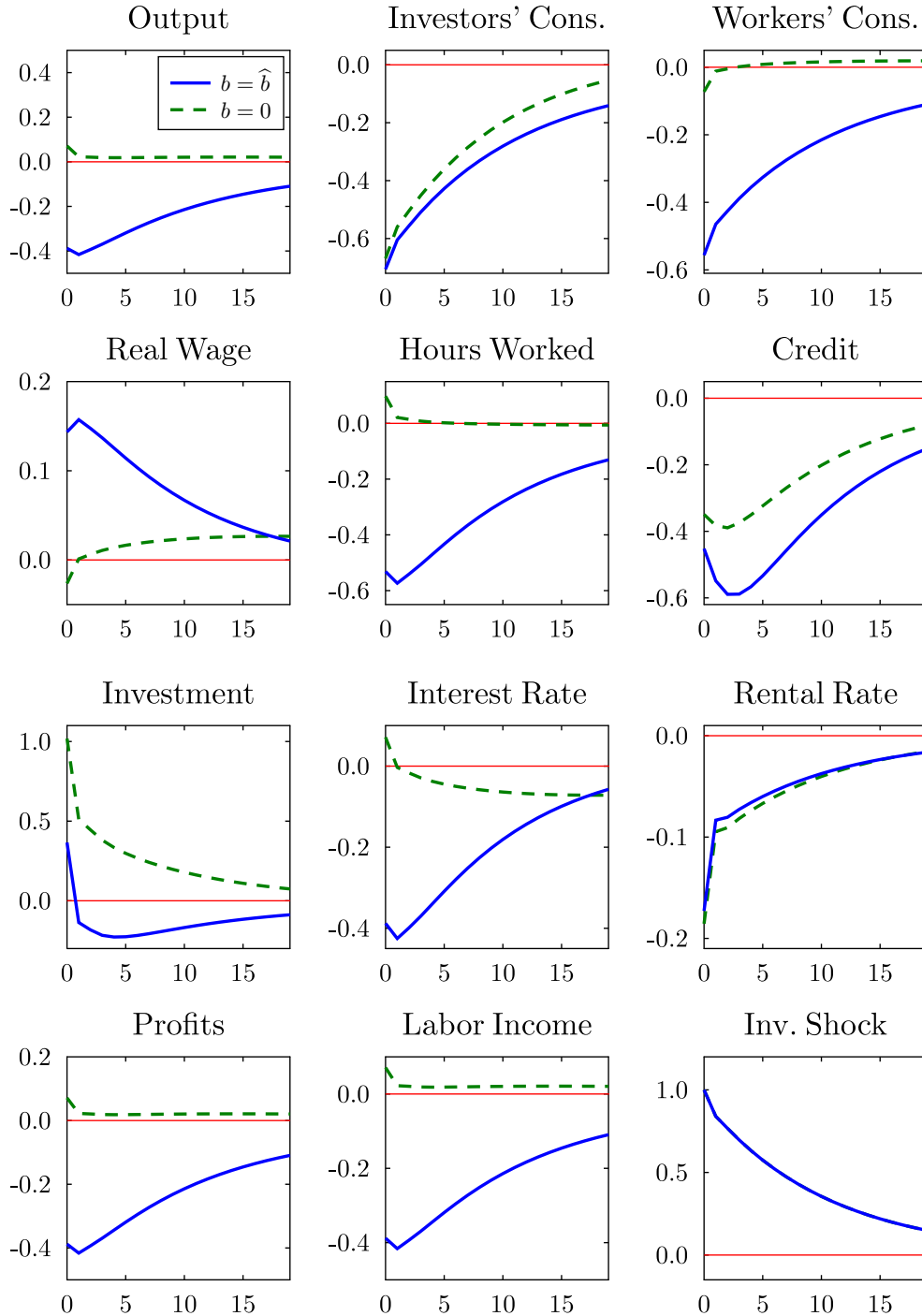
Figure 2.4 presents the model responses to the investment specific technology shock. Given the underlying AR(2)-structure, the response of the investment specific technology shock is more persistent than the respective shock responses described before. For the case of $b = \hat{b}$, investors increase investment on impact as the shock makes these expenditures more productive. However, in the following periods investment fall below its steady state value. On the other hand, investing households reduce their consumption level, a result often found in the standard represent agent framework (e.g. Fisher, 2006; Justiniano, Primiceri, and Tambalotti, 2010). By internalizing the investors con-

Figure 2.3: Impulse Responses to Price Markup Shock



Note: Responses are measured in percentage deviations from steady state. Horizontal axes measure time in quarters.

Figure 2.4: Impulse Responses to Investment Specific Technology Shock



Note: Responses are measured in percentage deviations from steady state. Horizontal axes measure time in quarters.

sumption decision, working households also decrease their consumption expenditures. This results in a reduced supply of hours worked and a falling demand for credit. The reduced labor supply induces an increase in the wage rate. As the fall in hours worked is stronger than the increase in investment, aggregate output and profits fall below their respective steady state values. Perfectly in line with the data, the models' responses to the investment specific technology shock lead to a negative correlation between credit and wages and positive co-movements between credit, hours worked, consumption, and output. The negative responses of almost all aggregate variables is supported by recent empirical evidence showing that investment-specific technology shocks have contractionary effects (Basu, Fernald, and Kimball, 2013).

The results change significantly when the consumption externality is switched off ($b = 0$). Working households now increase their labor supply and reduce their credit demand by a smaller amount, such that the reduction in their consumption expenditures is only marginally. Similarly, investors' consumption level drops less pronounced, also due to an increase in profits. Consequently, the rise in investment is more persistent and as both input factors increase also output goes up when the relative consumption motive is absent.

When we abstract from the consumption externality, the investment specific technology shock produces, in line with the data, a negative (positive) correlation between credit and wages (consumption) but negative co-movements between credit, output and hours worked which is in contrast to the empirical counterparts.

2.6 Conclusion

In this paper, we have set up a dynamic stochastic general equilibrium model that mimics the short-run dynamics of consumer credit for the period of the Great Moderation. The model consists of two different household types. Investors, who hold the economy's entire capital stock, own the firms and supply credit, and workers who make up the entire labor force and demand credit to finance their desired level of consumption. In

addition, we incorporate a *keeping up with the Riches* mechanism so that workers seek to minimize the difference between investors' and their own consumption level.

When estimating deep model parameter, we find a positive significant value for the workers' keeping-up parameter. Qualitatively, an income redistribution from labor to capital and an investment specific technology shock lead to model dynamics that are perfectly in line with the empirical evidence. More precisely, both shocks generate positive correlations of consumer credit with output, consumption, and labor, while there is a negative co-movement between consumer credit and the real wage. In contrast, a technology shock and a wage markup shock are not able to generate the positive correlations between consumer credit, output, and consumption. Complementary to micro evidence (Bertrand and Morse, 2013), we have provided macro-evidence on the link between income inequality, consumption externalities, and credit dynamics.

2.A Appendix

Table A2.1: Data Definitions and Sources

Variable	Definition	Source	Series ID (FRED database)
Consumer credit	Level of consumer credit held by households and nonprofit organizations	Board of Governors of the Federal Reserve System	HCCSDODNS
Output	Real output in the nonfarm business sector	U.S. Department of Labor: Bureau of Labor Statistics	OUTNFB
Hours worked	Hours of all persons in the nonfarm business sector	U.S. Department of Labor: Bureau of Labor Statistics	HOANBS
Real wage	Real compensation per hour in the nonfarm business sector	U.S. Department of Labor: Bureau of Labor Statistics	COMPRNFB
Labor productivity	Real output per hour of all persons in the nonfarm business sector	U.S. Department of Labor: Bureau of Labor Statistics	OPHNFB
Consumption	Real personal consumption expenditures	U.S. Department of Commerce: Bureau of Economic Analysis	PCECC96
Investment	Real gross private domestic investment	U.S. Department of Commerce: Bureau of Economic Analysis	GPDIC96
Capital	Net real capital stock of the total economy at 2005 prices (linear interpolation of annual values)	AMECO database of the European Commission	
Prices	Chain-type price index of personal consumption expenditures	U.S. Department of Commerce: Bureau of Economic Analysis	PCECTPI
Total factor productivity	Solow residual-based measure of technology corrected for labor and capital utilization, non-constant returns to scale and imperfect competition	Fernald (2012)	

3 Austerity and Private Debt¹²

ABSTRACT

This study provides empirical evidence that the costs of austerity crucially depend on the level of private indebtedness. In particular, fiscal consolidations lead to severe contractions when implemented in high private debt states. Contrary, fiscal consolidations have no significant effect on economic activity when private debt is low. These results are robust for alternative definitions of private debt overhang, the composition of fiscal consolidations, and controlling for the state of the business cycle and government debt overhang. I show that deterioration in household balance sheets is important to understand private debt-dependent effects of austerity.

Keywords: Fiscal consolidation, Private debt, Local Projection.

JEL Codes: C23, E32, E62.

3.1 Introduction

This study shows that the effects of fiscal consolidations crucially depend on the level of private indebtedness. More specifically, I find that austerity leads to severe contractions in periods of private debt overhang.¹³ In contrast, fiscal consolidations have no significant impact on economic activity when private debt is low.

The paper contributes to the literature as it tests for the validity of existing theoretical models which show that private indebtedness matters for the transmission of fiscal policy (for example Andrés, Boscá, and Ferri, 2015; Eggertsson and Krugman, 2012; Kaplan and Violante, 2014). In fact, I provide extensive empirical evidence that confirms predictions of theories pointing out the impact of fiscal policy interventions to be larger in periods of private debt overhang. My results help understanding

¹²A shortened version of this chapter is resubmitted to the *Journal of Money, Credit and Banking*.

¹³Private debt overhang describes periods when private debt-to-GDP ratios are above trend.

the dismal growth performances in southern European countries, which implemented large-scale fiscal consolidation programs while confronted with high private debt levels. To the best of my knowledge, this is the first empirical study investigating private debt-dependent effects of fiscal consolidations.

Recent contributions have pointed to the important role of private debt for the propagation and amplification of shocks and policy interventions. In their influential work, Mian and Sufi (2011, 2012) show that those US counties which experienced the largest increase in housing leverage before the financial crisis, suffered from more pronounced economic slack in the postcrisis period. The authors present evidence that deterioration in household balance sheets can explain the large drop in private demand and employment. Jordà, Schularick, and Taylor (2016b) find that more mortgage-intensive credit expansions tend to be followed by deeper recessions and slower recoveries, while this effect is not present for non-mortgage credit booms. Moreover, Jordà, Schularick, and Taylor (2016a) empirically investigate the linkage between private borrowing, public debt burdens, and financial instability and find that private credit booms, not excessive public borrowing or the level of public debt, are the main predictors of financial turmoil.

Concerning the interrelation between fiscal policy and private debt, Eggertsson and Krugman (2012), Kaplan and Violante (2014) and Andrés, Boscá, and Ferri (2015) demonstrate in theoretical models that the government spending multiplier increases with the level of private indebtedness. Within these models a significant share of households does not maximize lifetime utility due to borrowing constraints. Additionally, borrowing constrained households are characterized by a higher marginal propensity to consume out of income. Combined with price stickiness, Keynesian-type multipliers emerge if the share of these agents is large enough, which in turn depends on the level of private indebtedness.

Another strand of literature investigates state-dependent costs of fiscal consolidations (Born, Müller, and Pfeifer, 2015; Jordà and Taylor, 2016). None of these studies, however, allows the effects to differ according to the private debt level in the economy.

This seems surprising given the above mentioned evidence which suggests that the responses to economic innovations are amplified by private debt overhang. Against this background, I provide empirical evidence that the economic consequences of austerity are significantly affected by the level of private indebtedness.

To investigate the effects of fiscal consolidations depending on the state of the economy, I estimate state-dependent impulse responses to exogenous changes in the government budget deficit using local projections as invented by Jordà (2005). The advantages compared to vector autoregressions (VARs) are that local projections are more robust to model misspecification and offer a very convenient way to account for state dependence.¹⁴ Within the estimation approach, the state of the economy is allowed to vary according to the level of private debt overhang. High debt and low debt states are identified as periods when private debt-to-GDP ratios were respectively above and below trend. To identify fiscal consolidation periods, I use the narrative measure as proposed by Guajardo, Leigh, and Pescatori (2014). The baseline dataset of my analysis covers 12 OECD countries at an annual frequency for the period 1978-2008.

The estimation results show that the responses to fiscal consolidations significantly differ according to the level of private indebtedness. Specifically, the results reveal a significant and severe decline in private consumption and GDP in high debt states. Contrary, in low debt states, private consumption and GDP show a marginal and insignificant reduction. A one percent of GDP fiscal consolidation translates into a 2 percent lower GDP after five years when implemented in a period of private debt overhang. The drop in private consumption is even larger, resulting in a cumulative decline of more than 3 percent. The respective values for fiscal consolidations in low private debt states are 0.7 percent for GDP and 1.1 percent for consumption.

Concerning other important variables, I find that imports and the employment rate significantly decrease in high private debt states, whereas these series do not show any significant effect when private leverage is low. Monetary policy reacts to fiscal consolidations by reducing the real interest rate by a similar magnitude irrespective

¹⁴A more detailed discussion of advantages and disadvantages of the local projection method is given in the next section.

of the private debt state. Interestingly, the sovereign default risk and the government debt-to-GDP ratio increase significantly after consolidations implemented in a high private debt environment. This finding contradicts to the usual intention of austerity programs which lies in reducing the risk of sovereign default and/or reducing the government debt burden.¹⁵

My findings are robust for alternative definitions of debt overhang, different ways of identifying exogenous fiscal consolidation periods, and the composition of austerity programs. Moreover, I show that my baseline results prevail when extending the Guajardo, Leigh, and Pescatori (2014) narrative measure for the years 2010-2014. Thus, debt-dependent costs of fiscal consolidations are still present when explicitly taking into account the large-scale austerity programs implemented after the Global Financial Crises. In addition, the results prove to be robust when I condition on the state of the business cycle and government debt overhang.

Allowing the state of the business cycle to differ, I find that fiscal consolidations implemented in periods of high private debt induce economic activity to fall in recessions but also in booms. In expansions and recessions, austerity has no significant effect on the economy when private debt is below average. Similar results emerge when controlling for the government debt level. Independent of the government debt level, consolidations induce significant declines in economic activity when private leverage is high. In contrast, consolidations in low private debt states show insignificant effects irrespective of the public debt burden. To sum up, my findings suggest that the costs of austerity are mainly determined by the private debt level in the economy whereas the state of the business cycle and the level of public debt play only a minor role for the effectiveness of fiscal policy.

I highlight two additional results detecting changes in household balance sheets as a possible transmission channel through which my findings can be rationalized. First, by differentiating between household and corporate debt, I show that most of the results are driven by household leverage. While consolidations lead to a significant drop in

¹⁵Complementarily, Born, Müller, and Pfeifer (2015) show that austerity leads to an increase in the sovereign default premium in times of fiscal stress.

GDP when households are highly indebted, GDP does not react significantly when corporate debt is above average. Therefore, private debt-dependent effects of fiscal policy seem to be caused by households' and not firms' borrowing decisions. Second, house prices significantly decline when fiscal consolidations are implemented in high private debt states, whereas they basically do not show any effect in low private debt states. Falling house prices typically reduce the value of home equity households can use as collateral to borrow against.¹⁶

The closest related work to this study is the paper by Bernardini and Peersman (2015). They find that the government spending multiplier is considerably larger in periods of private debt overhang. However, my paper departs from their study in two important dimensions. First, while Bernardini and Peersman (2015) focus on non-linear effects of government spending, I estimate private debt-dependent responses to fiscal consolidations which are a combination of tax-based and spending-based adjustments. It seems reasonable to assume that the effects of austerity measures differ from standard fiscal spending shocks, because fiscal consolidations are typically implemented under special circumstances or because they are particularly large (Born, Müller, and Pfeifer, 2015). Moreover, it is unclear whether the effects of equally-sized expansion and tightening of fiscal policy should be symmetric, especially in the face of borrowing constraints. This argument is supported by recent empirical evidence showing that the government spending multiplier significantly differs between fiscal consolidations and fiscal expansions (Barnichon and Matthes, 2015; Riera-Crichton, Vegh, and Vuletin, 2015). In addition, I make use of the narrative consolidation measure to detect exogenous changes in fiscal policy. Second, my analysis is based on a panel dataset, whereas Bernardini and Peersman (2015) focus on the US economy. Thus, I provide multi-country evidence for private debt-dependent responses to fiscal policy.

The structure of the paper is organized as follows. In Section 2, the econometric method, database, and the identification of private debt states is described. Section 3

¹⁶As shown by Mian, Rao, and Sufi (2013), highly leveraged households have a higher marginal propensity to consume out of housing wealth such that, *ceteris paribus*, the aggregate drop in private demand to falling house prices increases with the level of private debt overhang in the economy.

presents results of the benchmark estimation. In addition, it is shown that my results are robust to an alternative identification approach, different ways of separating trend from cycle in private leverage, and when extending the narrative consolidation measure past 2009. In Section 4, I check whether the results depend on the composition of the fiscal consolidation. Moreover, I detect state-dependent effects of other relevant variables. In Section 5, I further control for two prominent state variables: the business cycle and government debt overhang. Section 6 presents evidence that indicates the importance of the household balance sheet for understanding private debt-dependent effects of fiscal consolidations. Finally, Section 7 concludes.

3.2 Econometric Method

To investigate the effects of fiscal consolidations depending on the state of the economy, I follow Auerbach and Gorodnichenko (2013), Ramey and Zubairy (2014) and Owyang, Ramey, and Zubairy (2013) in estimating state-dependent impulse responses to exogenous innovations in the government budget deficit using local projections as invented by Jordà (2005). Recently, this method has become a very popular tool to estimate non-linear effects. The main advantages compared to VARs are that local projections are more robust to model misspecifications and do not impose the implicit dynamic restrictions involved in VARs. Moreover, local projections offer a very convenient way to account for state dependence. However, the Jordà method does not uniformly dominate the standard VAR approach for calculating impulse responses. In particular, because it does not impose any restrictions that link the impulse responses across different horizons, the estimates are often erratic because of the loss of efficiency. Moreover, it sometimes display oscillations at longer horizons. For a more detailed discussion, I refer to Ramey and Zubairy (2014).

Let $Y_{i,t+h} - Y_{i,t-1}$ denote the cumulative response of a particular variable of interest from time $t - 1$ to $t + h$ to an exogenous change in the government budget deficit at time t , where i indexes the countries in my sample. I estimate a set of regressions of $Y_{i,t+h} - Y_{i,t-1}$ on shocks to the government budget deficit $D_{i,t}$ measured by the

narrative series as proposed by Guajardo, Leigh, and Pescatori (2014) and a set of control variables $X_{i,t}$:

$$\begin{aligned}
Y_{i,t+h} - Y_{i,t-1} = & I_{i,t-1} [\psi_{A,h}(L)X_{i,t-1} + \beta_{A,h}D_{i,t}] \\
& + (1 - I_{i,t-1}) [\psi_{B,h}(L)X_{i,t-1} + \beta_{B,h}D_{i,t}] + \alpha_{i,h} + \eta_{t,h} + \epsilon_{i,t+h}.
\end{aligned}
\tag{3.1}$$

Here, $\alpha_{i,h}$ are country-specific constants and $\eta_{t,h}$ captures time fixed effects to control for common macro shocks. $\epsilon_{i,t}$ denotes the error term which is assumed to have a zero mean and strictly positive variance. The dummy variable $I_{i,t}$ captures the state $\{A, B\}$ of the economy. $I_{i,t}$ takes the value of one when private debt is above a certain threshold and zero when it is below that threshold. Following the literature on state-dependent effects of fiscal policy (see for example Auerbach and Gorodnichenko, 2012; Ramey and Zubairy, 2014), I include a one-period lag of $I_{i,t}$ in the estimation to minimize the contemporaneous correlation between the shock series and changes in the indicator variable. L represents the lag operator. The collection of $\beta_{A,h}$ and $\beta_{B,h}$ coefficients directly provide the state-dependent responses of variable $Y_{i,t+h} - Y_{i,t-1}$ at time $t+h$ to the shock at time t . Given my specification, $\beta_{A,h}$ indicates the response of $Y_{i,t+h} - Y_{i,t-1}$ to the consolidation shock in high private debt states, whereas $\beta_{B,h}$ shows the effect in low private debt states.

In the following, all variables of interest are expressed in level log or level units. This stands in contrast to the approach used in Barro and Redlick (2011), Owyang, Ramey, and Zubairy (2013), and Ramey and Zubairy (2014) where the responses are scaled by GDP. However, given the facts that I use a much shorter sample compared to the aforementioned studies and that the consolidation shock $D_{i,t}$ is already scaled by GDP, it does not seem necessary to normalize the impulse responses by a measure of economic activity.

I prefer the specification of equation (3.1) to the propensity score matching method used in Jordà and Taylor (2016) because the former approach retains information about the size of fiscal consolidations, whereas the latter only allows the partition

of fiscal consolidations into a binary dummy variable 0/1 indicating periods of fiscal consolidation and periods of no consolidation. By retaining information about the magnitude of fiscal consolidations, I am able to directly measure the size of fiscal consolidation across different private debt states. Indeed, in Section 3.2.1, I show that my results are robust when controlling for anticipation effects in the narrative measure.

The dataset of my analysis is of annual frequency over the period 1978-2008 for a balanced sample of 12 OECD countries.¹⁷ The sample size of the panel is limited by the availability of the credit data used. In my baseline specification, the control variables included in $X_{i,t}$ are the absolute changes in the cyclically adjusted primary balance relative to GDP (CAPB), the log difference of real GDP and the log difference of real personal consumption expenditures.¹⁸ This choice closely mimics the VAR specification used in Guajardo, Leigh, and Pescatori (2014). $L = 1$ in all estimations, although the results are robust to varying the lag length.

To identify fiscal consolidation shocks, $D_{i,t}$, I use the narrative measure as proposed by Guajardo, Leigh, and Pescatori (2014). This measure is constructed by examining contemporaneous policy documents. The main advantage of identifying fiscal consolidations via the narrative measure compared to changes in the CAPB as suggested by Alesina and Ardagna (2010), is that the narrative measure is exogenous to current economic developments while changes in the CAPB are correlated to the business cycle. Guajardo, Leigh, and Pescatori (2014) show that there is a significant positive correlation between GDP forecast revisions and changes in the CAPB, whereas the null-hypothesis of no correlation between forecast revisions and the narrative measure cannot be rejected.

The definition of episodes of private debt overhang follows closely the approach by Bernardini and Peersman (2015). As an indicator for private debt, I use the private debt-to-GDP ratio, where data are taken from Schularick and Taylor (2012). Although

¹⁷The included countries are Australia, Canada, Germany, Denmark, Spain, France, United Kingdom, Italy, Japan, the Netherlands, Sweden and the United States. All data definitions and sources can be found in the appendix.

¹⁸The results are not affected when using *CAPB* in levels as control variable in the regressions. The appendix includes estimation results when controlling for CAPB instead of changes in the deficit variable.

the narrative consolidation measure is available for the period 1978-2009, Schularick and Taylor (2012) provide private debt data that just cover the years 1978-2008. To differentiate between high-debt and low-debt states, the debt-to-GDP ratios are filtered by country-specific smooth Hodrick-Prescott (HP) trends, where the smoothing parameter, λ , is set to 10,000. The relatively high smoothing parameter ensures that the filter removes even the lowest frequency variations in the private debt-to-GDP series. Indeed, the implementation of the Third Basel Accord (Basel III) involves the use of a similar credit gap indicator as used in my analysis (BIS, 2010). As shown by Borio (2014) and Drehmann, Borio, and Tsatsaronis (2012), the credit cycle is significantly longer and has a much greater amplitude than the standard business cycle. Therefore, Drehmann, Borio, and Tsatsaronis (2011) propose the use of an extremely smooth HP-trend to capture the low frequency of financial cycles. Given these considerations, applying an HP-filter with a smoothing parameter $\lambda = 10.000$ to construct the trending and cyclical component of private leverage seems appropriate for my analysis. High private debt states are defined as periods with positive deviations of the debt-to-GDP ratios from the trends, whereas low private debt states indicate periods when debt-to-GDP ratios were below its long-run trends. This procedure implies that out of the 372 periods included in the sample, 215 or 58% are detected as low private debt periods, while the remaining 157 episodes or 42% indicate periods of private debt overhang. In a separate exercise it is shown that the results are robust to two alternative definitions of high/low private debt states.

3.3 Results

3.3.1 Baseline

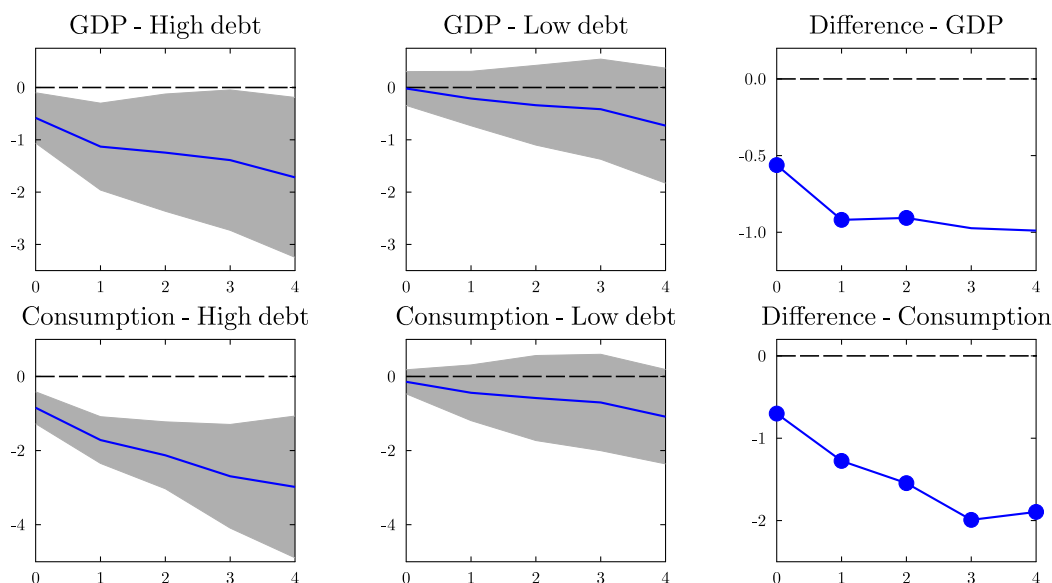
The main variables of interest, $Y_{i,t+h} - Y_{i,t-1}$, are the cumulative change in the log of real GDP and the cumulative change in the log of real personal consumption expenditures. Therefore, $\beta_{A,h}$ and $\beta_{B,h}$ directly estimate the state-dependent cumulative percentage change in the variables of interest in response to a fiscal consolidation shock.

Figure 3.1 presents the results of my baseline specification. It shows the cumulative effects on GDP and private consumption (solid lines) from year 0 to year 4 in response to a fiscal consolidation shock, where 0 indicates the year in which the shock occurs. Shaded areas indicate 90% confidence bands based on robust standard errors clustered by country. The respective responses are normalized so that the CAPB rises by 1% of GDP in year 0. The left column shows the cumulative responses to a fiscal consolidation implemented in a high private debt state, while the second column shows the respective changes to a fiscal consolidation undertaken in a low private debt state.

When private debt is below average, GDP shows a mild and insignificant reduction which accumulates to less than 1% four years after the fiscal consolidation was implemented. Contrary, fiscal consolidations undertaken when private leverage is high lead to a significant decline in GDP which accumulates to almost 2% at the end of the forecast horizon. A similar pattern can be observed for the respective consumption responses. Private consumption expenditures do not show a significant change in a low debt state. However, in a high private debt state consumption falls significantly such that expenditures are 3% lower after five years. The results indicate that a fiscal consolidation implemented when private debt is low leads to a small but insignificant reduction in economic activity, while fiscal consolidations in high private debt states induce a severe contraction in the economy.

Similar long-lasting, but non-permanent, negative effects of fiscal consolidations are found by Alesina and Ardagna (2010), Guajardo, Leigh, and Pescatori (2014), and Jordà and Taylor (2016). When estimating my baseline local projections for a longer horizon, all variables show a clear tendency to converge back to steady state values seven years after the fiscal consolidation was implemented. To rule out any instability concerns, I also estimated the model while including country-specific linear time trends. It turns out that the baseline results are not affected when controlling for a possible trending behavior in the endogenous variables. The results of both exercises can be found in the appendix.

Figure 3.1: Baseline Results



Note: The first two columns report cumulative changes (in per cent) in response to a shock of 1% of GDP to the cyclically-adjusted primary balance over $h = 0, 1, 2, 3, 4$ years. The shaded areas indicate 90% confidence bands based on robust standard errors clustered by country. The last column shows the estimated difference between high debt and low debt responses. Dots indicate statistically significant differences at the 90% level.

The right column of Figure 3.1 shows the respective differences $\beta_{A,h} - \beta_{B,h}$ for GDP and consumption at each period of the forecast horizon. Thus, a negative value indicates that the response in high debt states is lower than in low private debt states. The dots indicate statistical significance at the 90% level.

The response differences in GDP and private consumption are statistically significant for most of the periods. For GDP the differences are significant for 3 out of the 5 years, while they are significant for all 5 periods when inspecting the changes in private consumption. Complementarily to the first two columns of Figure 3.1, the latter findings indicate that the negative effects of austerity are significantly larger when the policy is implemented in a period of private debt overhang.¹⁹

As mentioned before, the Guajardo, Leigh, and Pescatori (2014) narrative consolidation measure covers the period 1978-2009. However, because the private debt data taken from Schularick and Taylor (2012) are just available for the period 1978-2008,

¹⁹The results are robust to changes in the sample. In the appendix it is shown that the estimates prevail when leaving out the years of the Global Financial Crises. In addition, it presents results indicating that my findings are not driven by any key country in the sample.

the baseline sample includes the years 1978-2008. Nevertheless, I am confident that the finding of private debt-dependent effects of fiscal consolidations is not affected by leaving out the year 2009 for three reasons. First, for the sample used the narrative measure does not identify any exogenous fiscal consolidation shock for the year 2009. Therefore, I do not expect the point estimates of my local projections to change significantly when adding observations of the final year 2009 to the sample. Second, I reestimate my baseline regressions using total credit data from the Bank of International Settlements (BIS). Contrary to the Schularick/Taylor series, they provide credit data for the year 2009. However, for my sample of interest the BIS-credit data only go back to 1980 so that I lose 12 observations compared to the (baseline) 1978-2008 sample. In the appendix it is shown that my findings prevail when using the BIS-credit data. Finally, as Section 3.3.2 shows, the result of private debt-dependent costs of fiscal consolidation is still present when using an extended version of the narrative measures such that the panel covers the years 1980-2014.

3.3.2 Robustness

Alternative Identification: Jordà and Taylor (2016) question the exogeneity of the narrative measure. They show that the Guajardo, Leigh, and Pescatori (2014) series has a predictable component. Therefore, my estimates could be biased when using the narrative measure as indicator for exogenous consolidation shocks.

To take account of possible anticipation effects, I combine the approach suggested by Jordà and Taylor (2016) with the forecast error-approach proposed by Auerbach and Gorodnichenko (2012).²⁰ The procedure consists of two steps. First, I regress the narrative consolidation measure, $D_{i,t}$, on a set of control variables which possibly include information that help predict the outcome variable. The residuals of this regression measure the unpredictable component of fiscal consolidations. In a second step, the residuals are used as proxy for exogenous austerity innovations in the estimation of equation (3.1).

²⁰Auerbach and Gorodnichenko (2012) use the unpredictable component of government spending as proxy for exogenous variations in fiscal expenditures.

Table 3.1: Alternative Identification of Fiscal Shock (effect in year $t = 1$)

Identification	GDP			Consumption		
	High Debt	Low Debt	Difference	High Debt	Low Debt	Difference
Baseline	-1.13*** (0.50)	-0.21 (0.32)	-0.92***	-1.72*** (0.38)	-0.44 (0.46)	-1.28***
Unpredictable component of D_{it}	-2.14*** (1.06)	0.05 (0.35)	-2.19***	-3.06*** (0.79)	-0.09 (0.53)	-2.97***

Notes: The table reports point estimates and robust standard errors clustered by country in parentheses. In each case the shocks are normalized so that the CAPB rises by 1% of GDP in year $t = 0$. *Significant at 16%; **significant at 10%; ***significant at 5%.

Motivated by the set of regressors chosen by Jordà and Taylor (2016), the vector of control variables in the first stage regression includes country and time fixed effects and a set of lagged macro variables (real GDP growth, real private consumption growth, change in government debt-to-GDP ratio, change in policy rate, CPI-inflation).

Table 3.1 presents the results when using the unpredictable component of $D_{i,t}$ as exogenous innovation and compares them to the benchmark estimation.²¹ As it turns out, my findings are robust to this alternative identification strategy. For both identification approaches, fiscal consolidations induce severe and significant reductions in GDP and private consumption when private debt is high, whereas in low private debt states both variables do not respond significantly. For both identifications, the GDP (consumption) response is estimated to be significantly lower when private leverage is high compared to low private leverage periods. This exercise shows that the finding of private debt-dependent effects of fiscal consolidation is robust to alternative ways of identifying fiscal consolidation episodes.

Alternative Debt States Definition: One possible concern with my baseline estimation could be that the results depend on the underlying definition of low and high private debt states. For this reason, I make use of two alternative ways to differentiate

²¹Following Guajardo, Leigh, and Pescatori (2014), all tables in the paper present the effects obtained two years after the fiscal consolidation was implemented (here, when $t = 1$). Moreover, the tables report whether the respective responses are statistically significant at the 5%, 10% and 16% level. The 16% level is chosen as lower threshold because of the relatively small sample size of the panel and because 16-84% confidence bands are widely used in the empirical macro literature (for example Bjørnland and Leitemo, 2009; Castelnuovo and Surico, 2010; Hofmann, Peersman, and Straub, 2012). For a general discussion on error bands for impulse responses see Sims and Zha (1999).

Table 3.2: Alternative Debt States Definition (effect in year $t = 1$)

Definition based on	GDP			Consumption		
	High Debt	Low Debt	Difference	High Debt	Low Debt	Difference
HP-filter (Baseline)	-1.13*** (0.50)	-0.21 (0.32)	-0.92***	-1.72*** (0.38)	-0.44 (0.46)	-1.28***
15-year MA	-0.64*** (0.31)	0.49 (0.71)	-1.14***	-1.37*** (0.31)	-0.75 (0.67)	-0.62***
Deviation from mean	-0.84** (0.47)	-0.31 (0.35)	-0.52*	-1.40*** (0.42)	-0.59 (0.43)	-0.80***

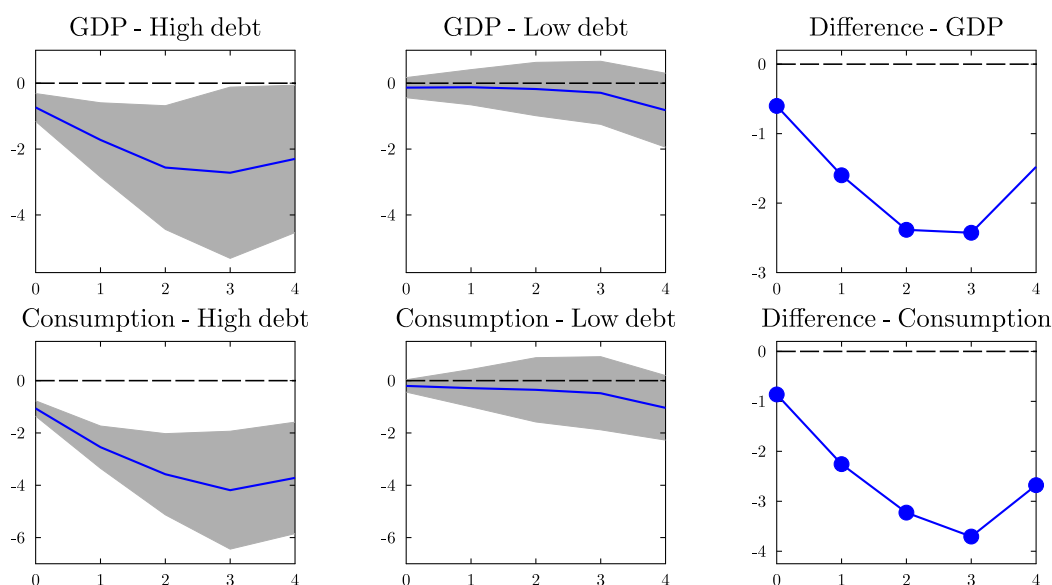
Notes: The table reports point estimates and robust standard errors clustered by country in parentheses. In each case the shocks are normalized so that the CAPB rises by 1% of GDP in year $t = 0$. *Significant at 16%; **significant at 10%; ***significant at 5%.

between high and low private debt periods. On the one hand, I calculate high (low) private debt episodes as periods in which the private debt-to-GDP ratio is above (below) its 15-year moving average. 15 years corresponds to the median length of financial cycles in industrialized countries (Borio, 2014). On the other hand, I follow Jordà, Schularick, and Taylor (2014) and define private debt states based on deviations from country-specific private leverage means. Whenever the change in the private debt-to-GDP ratio is above (below) its country-specific mean for two consecutive years, I define these episodes as high (low) private debt states.

As Table 3.2 shows, independent of the underlying debt state definition, I find strong evidence for private debt-dependent effects of fiscal consolidations. More precisely, GDP and private consumption decline significantly when private debt is high, whereas there is no significant response when private debt is low. Moreover, for all definitions, the respective GDP (consumption) response is estimated to be significantly lower when private debt is high compared to the corresponding low private debt one.

This exercise reveals that my findings do not rely on the specific way used to define low and high private debt states. The result of private debt-dependent costs of fiscal consolidations is robust to different definitions of private debt overhang. However, given the limited loss of observations compared to the other two definitions and its actual relevance in financial market policy (Basel III), in what follows I use the smooth

Figure 3.2: Extended Narrative Measure



Note: The first two columns report cumulative changes (in per cent) in response to a shock of 1% of GDP to the cyclically-adjusted primary balance over $h = 0, 1, 2, 3, 4$ years. The shaded areas indicate 90% confidence bands based on robust standard errors clustered by country. The last column shows the estimated difference between high debt and low debt responses. Dots indicate statistically significant differences at the 90% level.

HP-filter approach as the baseline method to separate trend from cyclical components in private leverage.

Extending the Narrative Measure: The baseline dataset covers the period 1978-2008, so it does not include the large-scale consolidation programs implemented by several countries in response to the significant increase in public debt levels following the deep economic downturn after the Global Financial Crises. To test whether my result of private debt-dependent effects of fiscal consolidations prevails when taking these austerity measures into account, I extend the Guajardo, Leigh, and Pescatori (2014) narrative series for the years 2010-2014.

In extending the dataset, I follow closely Dell'Erba, Mattina, and Roitman (2015) and Agca and Igan (2013) who construct a series of the consolidation measure for the years 2010 and 2011. The extension of the dataset is based on the following three OECD reports: *Restoring Public Finances, 2011*, *Restoring Public Finances, 2012 Update*, and *The State of Public Finances, 2015*. These reports outline the economic

situation, fiscal consolidation strategy and major consolidation measures for each of the OECD member countries. The country notes in each report lay out each government's rationale for pursuing fiscal adjustment and are used to identify consolidation periods that were motivated by a desire for deficit reduction. Table 2 of the appendix lists the identified consolidation periods for the years 2010-2014.²²

As the Schularick and Taylor (2012) loans series is just available until 2008, I make use of private credit data published by the Bank for International Settlements. To obtain private debt-to-GDP series, I divide the credit series by nominal GDP. Due to limited availability of the BIS credit data, the sample is now restricted to the period 1980-2014. As before, low/high private debt states are defined as deviations from a smooth HP-trend ($\lambda = 10.000$).

Figure 3.2 shows the impulse responses when using the extended narrative consolidation measure. Totally in line with the benchmark result, GDP and private consumption decrease significantly when private debt is high with slightly larger accumulated reductions compared to the baseline case. Contrary, GDP and private consumption do not respond significantly when private debt is below average. Additionally, the respective high debt responses are estimated to be significantly lower than the respective low debt ones for almost all periods. Thus, debt-dependent costs of fiscal consolidations are still present when explicitly taking into account the large-scale austerity programs implemented after the Global Financial Crises. Indeed, the results indicate that high private debt levels have amplified the negative effects of fiscal consolidations undertaken in the period 2010-2014.

²²I use the extension of the narrative measure as an additional robustness check and not as benchmark sample for two reasons. First, whereas the Guajardo, Leigh, and Pescatori (2014) measure is constructed by examining contemporaneous policy documents of various sources (IMF Reports, OECD Economic Surveys, Central Bank Reports, etc.), I rely mainly on the three OECD reports mentioned above. Second, it can be questioned to what extent the consolidations implemented between 2010 and 2014 can be treated as fully exogenous. Given the severity of the recession, the austerity programs undertaken in the aftermath of the Global Financial Crises could be related to the business cycle.

3.4 Extensions

In this section, I test whether the result of debt-dependent costs of austerity vary with the composition of the consolidation measure. Additionally, I show that the responses of other important macro variables also depend crucially on the private debt level when the consolidation is implemented.

3.4.1 Spending and Tax Based Consolidations

Guajardo, Leigh, and Pescatori (2014) and Alesina, Favero, and Giavazzi (2015) find that the costs of austerity differ with the composition of fiscal consolidations. Both studies show that tax-based consolidations lead to more severe contractions than spending-based adjustments. To allow the effects of consolidations to vary with its composition, I reestimate equation (3.1), where I make use of the composition definition stated by Guajardo, Leigh, and Pescatori (2014). The authors define fiscal policy changes as tax-based and spending-based if the budgetary impact of tax hikes and spending cuts, respectively, is greater than half the total impact.

Table 3.3 shows the estimates for spending-based and tax-based consolidations. Overall, the results coincide with the baseline estimation. Independent of the composition of the fiscal consolidation, GDP and private consumption do not change significantly when the austerity measure is implemented in a low private debt state. In contrast, GDP and private consumption are depressed significantly by tax-based and spending-based consolidations when private debt is high. In line with Alesina, Favero, and Giavazzi (2015) and Guajardo, Leigh, and Pescatori (2014), I find that tax-based consolidations have stronger effects on economic activity than spending-based adjustments. Nevertheless, my result of private debt-dependent costs of austerity is robust for the composition of fiscal consolidations.

3.4.2 Other Variables of Interest

So far, I have considered the private debt-depended responses of GDP and consumption to fiscal consolidations. However, it seems worth studying whether other important

Table 3.3: Spending Based vs. Tax Based (effect in year $t = 1$)

Composition	GDP		Consumption	
	High Debt	Low Debt	High Debt	Low Debt
Spending-based	-0.72** (0.43)	-0.21 (0.26)	-1.38*** (0.25)	-0.32 (0.39)
Tax-based	-5.20*** (1.82)	-0.59 (0.78)	-4.85*** (1.32)	-1.37 (1.07)

Notes: The table reports point estimates and robust standard errors clustered by country in parentheses. In each case the shocks are normalized so that the CAPB rises by 1% of GDP in year $t = 0$. *Significant at 16%; **significant at 10%; ***significant at 5%.

macro variables also react differently to fiscal consolidations in high and low private debt periods. In the following, I check for divergent responses in other components of GDP: private investment, imports, and exports. Moreover, I test whether the effects on the labor market, measured through the employment rate, differ as well. It is shown that the central bank reduces its main policy rate by a similar magnitude, irrespective of the private debt state. Finally, I study how the sovereign default risk, indicated by the institutional investor ratings index (IIR) and the government debt-to-GDP ratio response to consolidations in both private debt states. At each horizon, I project these variables on fiscal consolidations and include their respective lags in the control vector $X_{i,t}$. While investment, imports, and exports enter the estimation in log differences, the employment rate, interest rate, IIR, and the government debt-to-GDP ratio are considered in absolute changes.

Figure 3.3 presents the responses of investment, imports, and exports. Private investment increases slightly when the consolidation is undertaken in a period of low private debt. However, this increase is not statistically significant. In high private debt states, investment decreases significantly by more than 2% in the first two years. Afterwards, the effect becomes insignificant as well. The mostly insignificant investment response relates to the empirical evidence presented by Mian, Sufi, and Verner (2015). They show that rises in household debt are closely tied to consumption and less related to business investment. Additionally, it can be interpreted as a first indicator that households', not firms', borrowing decisions are mainly responsible for private debt-dependent effects of austerity. However, below I will elaborate in more detail

on the household balance sheet as a possible transmission channel to rationalize my findings.

Divergent responses can also be observed for imports. Imports decrease slightly but insignificantly in low private debt states. In contrast, imports are more than 5% lower after 5 years when the consolidation is undertaken in a high private debt period. The difference in the respective import responses is significant for all periods.

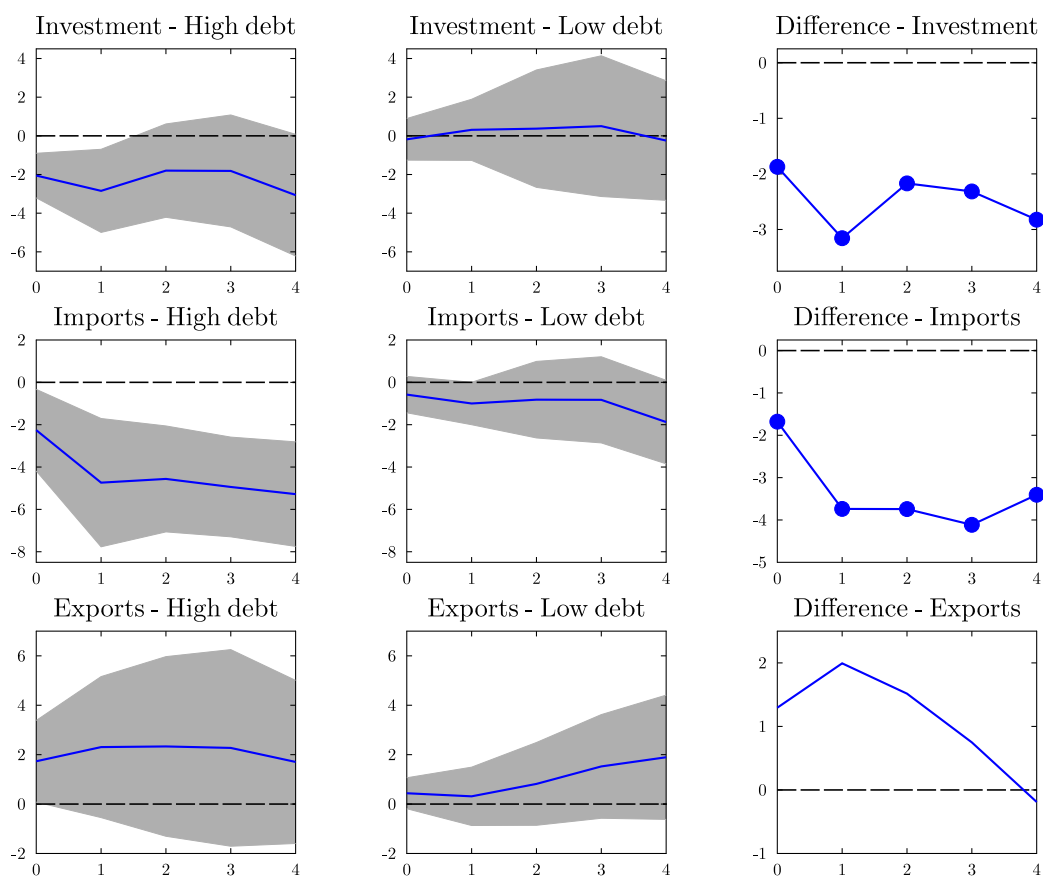
In both debt states exports increase substantially. However, the respective responses are not statistically different from zero for most of the periods. As exports react rather similar in low and high debt states, the response difference is not statistically significant.²³

Figure 3.4 shows the results for the employment rate, interest rate, IIR, and government debt. The employment rate increases steadily when private debt is below average. Consolidations in high private debt states lead to a significant decline in the employment rate. The accumulated loss after four years is 1.5 percentage points. Additionally, as the right column shows, the employment rate response in high private debt states is significantly lower than the respective one in low private debt states. These findings indicate that the severe real costs of fiscal consolidations implemented when private debt is high also translate into a deterioration in the labor market. This relation is also captured by the theoretical set-up by Andrés, Boscá, and Ferri (2015). In their model, the improvement in the labor market to a government spending shock depends positively on the equilibrium level of household debt.

Private debt-dependent responses to fiscal consolidations could be explained by a different reaction of the monetary authority to austerity in low and high debt states. When the central bank reduces (increases) its interest rate by less (more) when austerity is realized in a high leverage period compared to a low debt state, then the more severe downturn could be caused by a debt-dependent interest rate change. Indeed, as the second row of Figure 3.4 demonstrates, this hypothesis is not supported by the

²³Taking the effects on imports and exports together, in an additional exercise, I found that the current account significantly increases in high private debt states, while it stays almost unchanged when private debt is low.

Figure 3.3: Investment, Imports, Exports



Note: The first two columns report cumulative changes (in per cent) in response to a shock of 1% of GDP to the cyclically-adjusted primary balance over $h = 0, 1, 2, 3, 4$ years. The shaded areas indicate 90% confidence bands based on robust standard errors clustered by country. The last column shows the estimated difference between high debt and low debt responses. Dots indicate statistically significant differences at the 90% level.

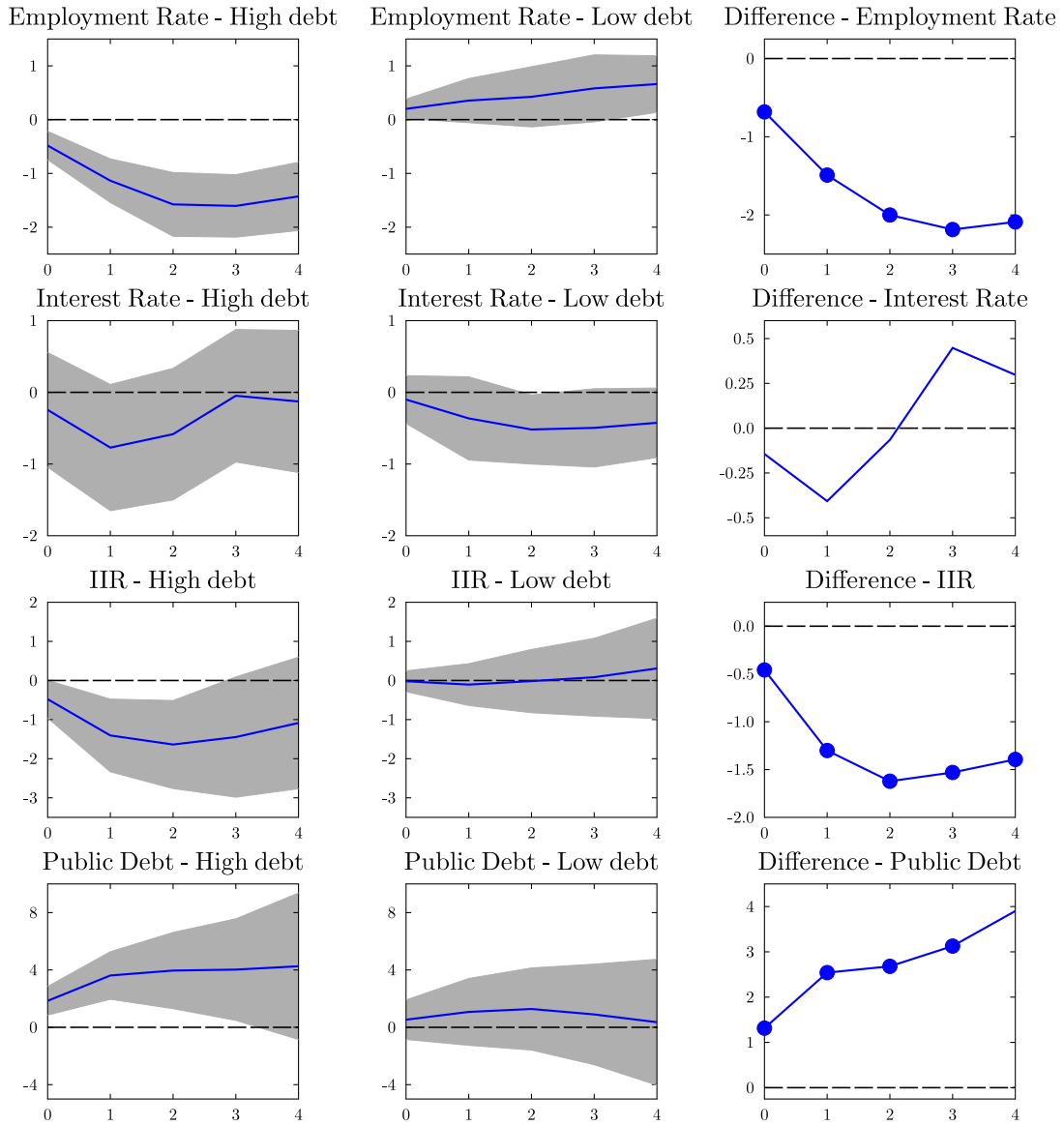
data. The central bank reduces the interest rate by a similar magnitude irrespective of the private debt state. Overall, both interest rate responses are insignificant for almost all periods indicating a rather conservative expansionary monetary policy in reaction to fiscal consolidations. Not surprisingly, the response difference is statistically insignificant for all years of the forecast horizon.

The IIR is based on assessments of sovereign default risk by private sector analysts on a scale of zero to 100, with a rating of 100 assigned to the lowest perceived sovereign default probability. As the third row of Figure 3.4 shows, the index falls when consolidations are implemented in a high private debt state, implying a higher probability of sovereign default. Significant reductions in the IIR are visible up to three years after the consolidation. Interestingly, even in low debt states the IIR does not increase but mainly stays unchanged 4 years after the implementation took place. In all of the five periods, the high debt IIR response is significantly lower than the low debt IIR response.

Finally, I look at how the government debt-to-GDP ratio is affected by fiscal consolidations in both private debt states. In high private debt states, the public debt burden shows a persistent and significant increase which accumulates to a rise of more than 4 percentage points at the end of the forecast horizon. In contrast, the government debt-to-GDP ratio does not respond significantly when private leverage is below average. In addition, the high debt response is estimated to be significantly larger than the respective low debt response in four out of the five periods considered. In contrast to reducing public debt burdens which is one of the main goals of fiscal consolidations, public debt burdens even increase when private debt is high. Together with the effects on the sovereign default probability, this finding indicates that austerity in high private debt states is not only associated with high costs for the private sector but also with a worsening of government finances.

To summarize, besides GDP and consumption, also imports, the employment rate, the sovereign default risk and, the government debt-to-GDP ratio react differently to fiscal consolidations depending on the private debt level in the economy.

Figure 3.4: Employment, Interest Rate, Investors' Confidence, Public Debt



Note: The first two columns report cumulative changes (in percentage points) in response to a shock of 1% of GDP to the cyclically-adjusted primary balance over $h = 0, 1, 2, 3, 4$ years. The shaded areas indicate 90% confidence band based on robust standard errors clustered by country. The last column shows the estimated difference between high debt and low debt responses. Dots indicate statistically significant differences at the 90% level.

3.5 Additional State Variables

In this section it is demonstrated that the result of private debt-dependent effects of austerity still prevails when I further condition on two other prominent state variables: the state of the business cycle and government debt overhang.

3.5.1 Booms and Recessions

Jordà and Taylor (2016) show that the costs of fiscal consolidations differ according to the state of the business cycle. They find that austerity leads to a significant drop in economic activity when implemented in recessions while there is no significant effect when consolidations are undertaken in a boom. Additionally, Auerbach and Gorodnichenko (2012) present empirical evidence that the government spending multiplier is larger in periods of economic slack. Contrary, Ramey and Zubairy (2014) do not find significant differences between spending multipliers in good and bad times. To check whether my findings are sensitive to the state of the business cycle, I further condition equation (3.1) on expansionary and recessionary states. Thereby, I make use of three common approaches to differentiate between expansionary and recessionary periods. As a benchmark case, I use the recession dates published by the OECD. Second, similar to Jordà and Taylor (2016), I calculate the cyclical component of GDP measured as deviations from (country-specific) HP trends with a smoothing parameter $\lambda = 6.25$ as suggested by Ravn and Uhlig (2002). Positive deviations from the trend are defined as booms and negative deviations as recessions. Third, following the approach proposed by Auerbach and Gorodnichenko (2012), I construct (country-specific) four-year moving averages of real GDP growth, and classify periods as expansions (recessions) whenever the actual growth rate is above (below) the moving average rate.

I reestimate equation (3.1) separately for low and high private debt states based on the following equation:

$$\begin{aligned}
Y_{i,t+h} - Y_{i,t-1} &= I_{C,i,t-1} [\psi_{C,h}(L)X_{i,t-1} + \beta_{C,h}D_{i,t}] \\
&+ I_{D,i,t-1} [\psi_{D,h}(L)X_{i,t-1} + \beta_{D,h}D_{i,t}] \\
&+ I_{E,i,t-1} [\psi_{E,h}(L)X_{i,t-1} + \beta_{E,h}D_{i,t}] + \alpha_{i,h} + \eta_{t,h} + \epsilon_{i,t+h}.
\end{aligned} \tag{3.2}$$

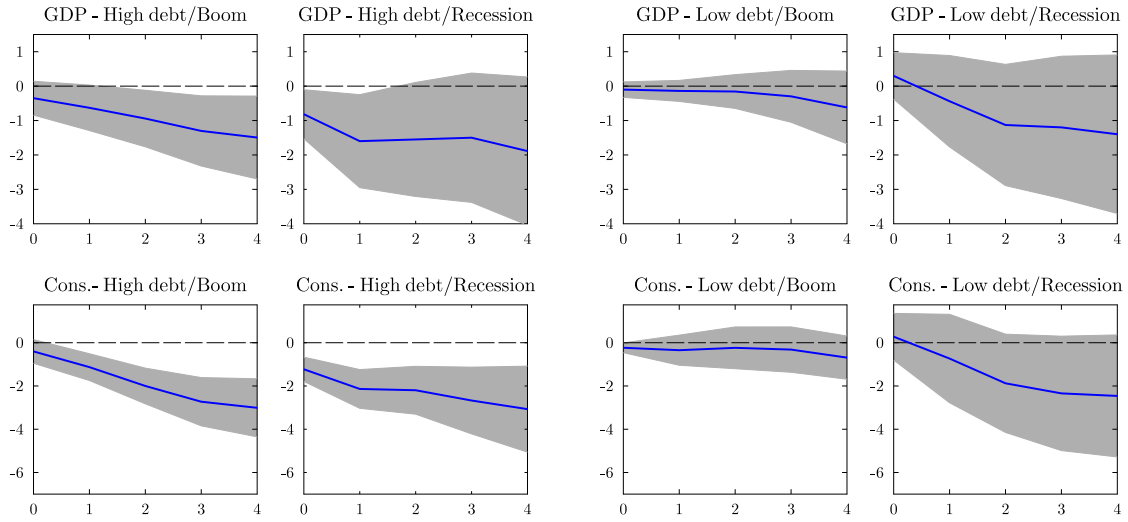
$I_{C,i,t}$ and $I_{D,i,t}$ now indicate the state of the business cycle of the respective private debt states. In the estimation for high private debt states, $I_{C,i,t}$ measures periods of high private debt that coincide with periods of economic contractions whereas $I_{D,i,t}$ indicates periods of high private debt that are also characterized by economic expansions. $I_{E,i,t}$ is then a dummy variable for being in the opposing private debt state (low private debt), irrespective of the state of the business cycle. Analogously, in the estimation for low private debt states, $I_{C,i,t}$ ($I_{D,i,t}$) measures periods of low private debt that coincide with periods of economic contractions (expansions) and $I_{E,i,t}$ indicates periods of high private debt. $\beta_{C,h}$ and $\beta_{D,h}$ then provide the state-dependent responses for both debt states in recessions and booms, respectively.

Figure 3.5 shows the results based on the OECD business cycle classification, whereas Table 3.4 reports the effects when using the two other classification strategies. Independent of the business cycle classification applied, when private debt is high, GDP and consumption decline significantly in recessionary but also in expansionary periods. In contrast, in low private debt states the effects of fiscal consolidations are not significant neither in booms nor in recessions. Moreover, the size of the respective point estimates in both business cycle states is fairly similar indicating that business cycle-dependent effects of fiscal consolidations disappear when controlling for private leverage in the economy.

3.5.2 Government Debt

In addition to the state of the business cycle, previous literature found that the effects of fiscal policy vary with the level of public debt in the economy. Perotti (1999) shows that an increase in government consumption leads to higher private consumption

Figure 3.5: Controlling for State of the Business Cycle



Note: Cumulative changes (in per cent) in response to a shock of 1% of GDP to the cyclically-adjusted primary balance over $h = 0, 1, 2, 3, 4$ years. The shaded areas indicate 90% confidence bands based on robust standard errors clustered by country.

expenditures when government debt is low, whereas consumption declines when public debt-to-GDP levels are high. Similar, Ilzetzki, Mendoza, and Végh (2013) provide evidence that the government spending multiplier negatively depends on the public debt level.

To check whether the result of private debt-dependent costs of fiscal consolidations still holds when controlling for the public debt level, I reestimate equation (3.2) where $I_{C,i,t}$ and $I_{D,i,t}$ now indicate the respective government debt levels in the periods of both private debt states. In the estimation for high private debt states, $I_{C,i,t}$ measures periods of high private debt that coincide with periods of low government debt, whereas $I_{D,i,t}$ indicates periods of high private debt that are also characterized by high public debt levels. $I_{E,i,t}$ is then a dummy variable for being in the opposing private debt state (low private debt), irrespective of the government debt level.²⁴ Periods of high (low) public debt are defined as positive (negative) deviations of the government debt-to-GDP ratio from a country-specific smooth HP trend ($\lambda = 10,000$).

²⁴Analogously, in the estimation for low private debt states, $I_{C,i,t}$ ($I_{D,i,t}$) measures periods of low private debt that coincide with periods of low (high) public debt burdens and $I_{E,i,t}$ indicates periods of high private debt.

Table 3.4: Alternative Business Cycle Classification (effect in year $t = 1$)

Classification based on Detrended GDP	Boom		Recession	
	High Debt	Low Debt	High Debt	Low Debt
GDP	-1.22* (0.79)	-0.07 (0.33)	-0.97** (0.49)	-0.42 (0.41)
Consumption	-1.75*** (0.70)	-0.28 (0.38)	-1.59*** (0.39)	-0.56 (0.61)
Classification based on MA GDP growth	Boom		Recession	
	High Debt	Low Debt	High Debt	Low Debt
GDP	-1.08* (0.79)	-0.11 (0.21)	-0.98** (0.58)	-0.47 (0.65)
Consumption	-1.56*** (0.73)	-0.34 (0.43)	-1.60*** (0.52)	-0.55 (0.79)

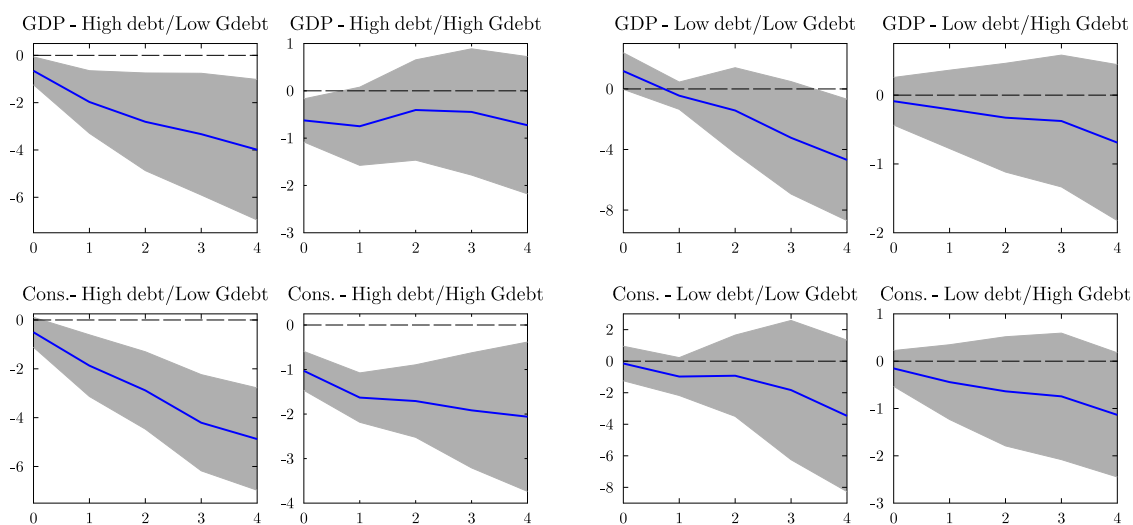
Notes: The table reports point estimates and robust standard errors clustered by country in parentheses. In each case the shocks are normalized so that the CAPB rises by 1% of GDP in year $t = 0$. *Significant at 16%; **significant at 10%; ***significant at 5%.

Figure 3.6 presents the cumulative responses for both private debt states when controlling for the public debt burden. GDP and private consumption decline significantly irrespective of the public debt level when private debt is high. In line with the findings by Perotti (1999) and Ilzetki, Mendoza, and Végh (2013), the effects are larger in periods of low public debt. When government debt is low, GDP (consumption) is 3.9% (4.8%) lower four years after the implementation. In high government debt states, the accumulated loss is 0.7% for GDP and 2.1% for consumption.

Turning to the low private debt responses, I find insignificant effects for periods with high public debt burdens. When government debt is low, GDP shows a significant response only in the last period of the forecast horizon, whereas consumption does not react significantly in all periods. In accordance to the respective high private debt responses, the point estimates for GDP and consumption are larger when the government debt level is low.

To sum up, the last two exercises demonstrate that fiscal consolidations implemented in high private debt states are always a drag on private economic activity, irrespective of the state of the business cycle or the government debt level. In contrast, austerity

Figure 3.6: Controlling for Government Debt Level



Note: Cumulative changes (in per cent) in response to a shock of 1% of GDP to the cyclically-adjusted primary balance over $h = 0, 1, 2, 3, 4$ years. The shaded areas indicate 90% confidence bands based on robust standard errors clustered by country.

measures undertaken in low private debt periods do not have a significant effect on the economy in booms and recessions, when government debt is high or low. This result gives rise to the interpretation that effectiveness of fiscal policy does not vary with the business cycle or the public debt burden but rather with the level of private leverage. Whether this reasoning also contributes to the controversial debate on state-dependent government spending multipliers (see for example Auerbach and Gorodnichenko, 2012; Ramey and Zubairy, 2014) could be an interesting agenda of future research.

3.6 Household Balance Sheet

What is the underlying transmission channel through which my results can be rationalized? In the following, I present evidence indicating that deterioration in household balance sheets as proposed by Mian and Sufi (2011, 2012) is of central importance for understanding private debt-dependent responses to fiscal consolidations. They stress that the large drop in private demand during the Great Recession was mainly caused by a worsening in housing net worth of highly leveraged households. Moreover, U.S. counties with a larger decline in housing net worth were found to experience a larger de-

cline in employment. In a recent paper, Mian, Sufi, and Verner (2015) empirically show that an increase in private debt is associated with lower output growth in the future. This result only holds for increases in household debt, while for rises in corporate debt the authors do not find significant future output effects. In a theoretical framework, Andrés, Boscá, and Ferri (2015) show that the spending multiplier increases with the level of households' indebtedness. Their model economy is populated by two types of households, lenders and borrowers. Borrowing households face a collateral constraint which limits the maximum loans that an individual can get to a fraction of the liquidation value of the amount of housing held by the household, the loan-to-value ratio. By assuming that the collateral constraint holds with equality in equilibrium, it can be shown that borrowing households discount the future more heavily than lending households. This model feature is backed by the empirical finding that indebted households have a higher marginal propensity to consume out of housing wealth (Mian, Rao, and Sufi, 2013). In a simulation exercise, Andrés, Boscá, and Ferri (2015) show that the size of the spending multiplier positively depends on the share of borrowers in the economy and the loan-to-value ratio, which in turn depend on the level of indebtedness. Taken together, all these studies find that a high level of household indebtedness amplifies the effects to economic shocks.

The central determinant of housing net worth are real estate prices. Mian and Sufi (2011, 2012) demonstrate that changes in house prices crucially affect private consumption expenditures. Falling house prices led to a deterioration in households balance sheets which, through the housing net worth channel, resulted in the large reduction in economic activity observed during the Great Recession. Andrés, Boscá, and Ferri (2015) model house prices as one variable of the liquidation value households can use as collateral to borrow against.

Given these considerations, my results are tested in two additional dimensions. First, I split private debt into household debt and corporate debt and check whether my findings depend on the specific type of private leverage. Second, I show how house prices respond to fiscal consolidations in high and low private debt states.

Table 3.5: Household Debt vs. Corporate Debt (effect in year $t = 1$)

Private debt type	Household debt			Corporate Debt		
	High Debt	Low Debt	Difference	High Debt	Low Debt	Difference
GDP	-0.89** (0.53)	-0.01 (0.57)	-0.87**	-0.69 (0.56)	-0.45 (0.46)	-0.25

Notes: The table reports point estimates and robust standard errors clustered by country in parentheses. In each case the shocks are normalized so that the CAPB rises by 1% of GDP in year $t = 0$. *Significant at 16%; **significant at 10%; ***significant at 5%.

Table 3.5 presents the different GDP responses in low/high corporate debt and low/high household debt states. Equation (3.1) is separately estimated for both types of private debt. Series on corporate debt and household debt are taken from the Bank for International Settlements, where, due to data limitations, the panel is now restricted to the period 1980-2008 and the countries Australia, Canada, Germany, Spain, France, United Kingdom, Italy, Japan, Sweden and the United States. To obtain private debt-to-GDP series, I divide the respective debt series by nominal GDP. As before, low/high corporate debt and household debt periods are identified as deviations from a smooth trend (HP-filter with $\lambda = 10,000$).

It turns out that my major finding of private debt-dependent effects of fiscal consolidations is mainly driven by households' leveraging position and not corporate debt overhang. The fall in GDP in response to austerity is not significant when corporate debt is high. In contrast, GDP declines significantly when private households are highly leveraged. Although the effect in high corporate debt states is somewhat larger than in low corporate debt states, the difference between both responses is statistically insignificant. A different picture emerges for household debt. The response difference between high and low household debt states is estimated to be statistically significant. In line with the findings by Mian and Sufi (2011, 2012) and Andrés, Boscá, and Ferri (2015), the results in Table 3.5 point to the important role of household leveraging for the economic dynamics to fiscal interventions. Corporate debt levels do not seem to be responsible for understanding private debt-dependent effects of fiscal policy.

Given the prominent role of households' leveraging position for understanding my results, it seems natural to investigate how the central driver of housing wealth, house

prices, react to fiscal consolidations in low/high private debt periods. As mentioned earlier, house prices are one key ingredient of households' optimal consumption decision. Falling house prices reduce the home equity value that serves as collateral to borrow against, which ultimately results in lower consumption expenditures by constrained agents (Mian and Sufi, 2011, 2012). To test whether this transmission channel also applies to my findings, Figure 3.7 shows the response of house prices to fiscal consolidations implemented in low and high private debt states. House price data are taken from the Federal Reserve Bank of Dallas (Mack and Martínez-García, 2011). At each horizon, house prices are projected on fiscal consolidations and their respective lag is included in the vector of control variables $X_{i,t}$. House prices enter the estimation in log differences.

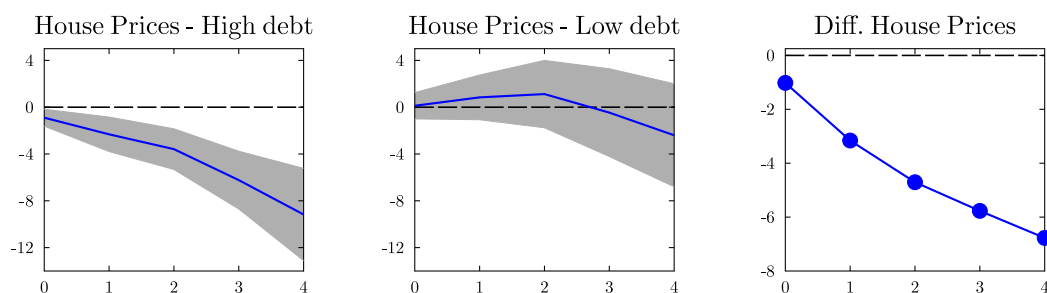
Figure 3.7 shows that the response of house prices crucially depends on the private debt level when the fiscal consolidation is undertaken. House prices do not react significantly when private leverage is low. However, in a high private debt state house prices significantly fall with a accumulated decline of almost 10% after five years. As the last column of Figure 3.7 demonstrates, the difference between the respective responses is statistically significant for all five periods.

Although causal interpretations should be taken cautiously, the evidence shown in this section indicates that private debt-dependent costs of fiscal consolidations can be rationalized through deterioration in household balance sheets. Theories should, therefore, elaborate on the housing net worth channel (Mian and Sufi, 2011, 2012) when studying the consequences of fiscal policy interventions.

3.7 Conclusion

Motivated by recent theoretical contributions that show the effects of fiscal policy to be larger in periods of high private leverage (see for example Andrés, Boscá, and Ferri, 2015; Eggertsson and Krugman, 2012; Kaplan and Violante, 2014), this paper has shown that the level of private indebtedness significantly determines the costs of fiscal

Figure 3.7: House Prices



Note: Cumulative changes (in per cent) in response to a shock of 1% of GDP to the cyclically-adjusted primary balance over $h = 0, 1, 2, 3, 4$ years. The shaded areas indicate 90% confidence bands based on robust standard errors clustered by country.

consolidations. Based on a panel of 12 OECD countries, I use local projection methods, which allow responses to differ between low debt and high debt states.

I find that austerity implemented in a low private debt state does not induce significant changes in GDP and private consumption. In contrast, fiscal consolidations lead to severe contractions in GDP and private consumption when private debt is high. This result is robust to different ways of identifying fiscal consolidations, alternative definitions of low/high private debt states, the composition of fiscal consolidations, controlling for the state of the business cycle and government debt overhang. In addition, the finding of private debt-dependent costs of fiscal consolidations is still present when extending the sample such that it includes large-scale austerity programs implemented in the period 2010-2014. Imports and employment fall significantly when private leverage is high, while they do not show any significant effect when private debt is low. Moreover, in high private debt states, consolidations lead to a persistent increase in the government debt-to-GDP ratio which contradicts with one of the main goals of fiscal austerity that lies in reducing public debt burdens.

Two additional findings highlight the importance of the housing net worth channel (Mian and Sufi, 2011, 2012) for understanding my results. First, the private debt-dependent responses to fiscal consolidations are mainly driven by household debt and not corporate debt. Second, I show that house prices significantly decline when consolidations are implemented in a period of private debt overhang. Both of these latter

observations indicate that deterioration in household balance sheets represents a possible channel through which my results can be explained.

My findings reveal important implications. They confirm predictions of theoretical models as the ones by Eggertsson and Krugman (2012), Kaplan and Violante (2014) and Andrés, Boscá, and Ferri (2015), which point out the impact of fiscal policy interventions to be larger in periods of private debt overhang. Moreover, high private debt levels in Southern European countries may have amplified the negative effects of large-scale fiscal consolidations. Contrary to its objective of improving public finances, austerity measures could have even increased solvency problems. More generally speaking, the level of private debt and especially of household debt seems to matter for the effects of fiscal policy.

3.A Appendix

This appendix includes all data definitions and sources and reports the results of additional estimation results and robustness checks mentioned in the text.

Table A3.1 presents all data definitions and sources.

Table A3.2 reports the identified narrative fiscal shocks for the period 2010-2014.

Table A3.3 shows that my results are not affected when using $CAPB_t$ in levels as control variable in the regressions. The baseline results still hold when controlling for $CAPB_t$ instead of $\Delta CAPB_t$.

Table A3.4 demonstrates that private debt-dependent effects of fiscal consolidations also emerge when using credit data from the Bank for International Settlements (BIS). The local projections using the BIS-credit data are based on the years 1980-2009.

Table A3.5 shows the effects when controlling for country-specific time trends in my baseline specification. It turns out, that my main finding is not affected when allowing for a possible trending behavior in the endogenous variables.

To rule out that my results are driven by the Global Financial Crises, Table A3.6 presents the results when considering the 1978-2006 sample. My estimates are robust to leaving out the Crises years.

To assess how important any individual country is for the results, I reestimate the local projections, while dropping one country at a time from the sample. As Table A3.7 indicates, the results are comparable to the baseline in each case.

Figure A3.1 presents results when estimating the baseline regressions for a longer horizon. All variables show the tendency to converge back to steady state seven years after the consolidation was implemented. This gives rise to the interpretation that fiscal consolidations have long-lasting, but non-permanent negative effects. Together with Table A3.5 this finding indicates that my findings are not driven by unstable impulse responses.

The baseline sample covers the period 1978-2008 and the countries Australia, Canada, Germany, Denmark, Spain, France, United Kingdom, Italy, Japan, the Netherlands, Sweden and the United States.

Table A3.1: Data Definitions and Sources

Variable	Definition	Source
GDP, real	Gross domestic product, constant prices, OECD base year	OECD
GDP, nominal	Gross domestic product, current prices, current PPPs, in US Dollar	OECD
Consumption	Final consumption expenditures, households and non-profit institutions serving households, constant prices, OECD base year	OECD
CAPB	Cyclically-adjusted primary balance	Alesina and Ardagna (2010)
Private debt to GDP	End-of-year amount of outstanding domestic currency lending by domestic banks to domestic households and nonfinancial corporations (excluding lending within the financial system) to GDP	Schularick and Taylor (2012)
Fiscal consolidation	Changes in fiscal policy motivated by a desire to reduce the budget deficit and not by responding to prospective economic conditions	Guajardo, Leigh, and Pescatori (2014)
Investment	Gross fixed capital formation, constant prices, OECD base year	OECD
Imports	Imports of goods and services, constant prices, OECD base year	OECD
Exports	Exports of goods and services, constant prices, OECD base year	OECD
Employment rate	Civilian employment as % population (15-64 years old)	OECD
Interest rate	Main central bank policy interest rate	Guajardo, Leigh, and Pescatori (2014)
Institutional Investors Rating Index	Assessments of sovereign risk by private sector analysts on a scale to 100, with a rating of 100 assigned to the lowest perceived sovereign default probability	Guajardo, Leigh, and Pescatori (2014)
Household debt	End-of-year credit to households and NPISHs from all sectors, market value, in US Dollar, adjusted for breaks	Bank for International Settlements; sample restricted to 1980-2008, no data for Denmark and Netherlands
Corporate debt	End-of-year credit to non-financial corporations from all sectors, market value, in US Dollar, adjusted for breaks	Bank for International Settlements; sample restricted to 1980-2008, no data for Denmark and Netherlands
Total credit to private sector	End-of-year credit to private non-financial sector from all sectors, market value, in US Dollar, Adjusted for breaks	Bank for International Settlements; sample restricted to 1980-2014
House prices	Real house prices index (four-quarter average)	Federal Reserve Bank of Dallas (Mack and Martínez-García, 2011)
Public debt to GDP	Face value of total general government debt outstanding to GDP	Jordà, Schularick, and Taylor (2016a)
OECD recession indicator	OECD based recession indicator from the peak through the trough	OECD

In extending the narrative consolidation measure, I closely follow Dell’ Erba, Mat-
tina, and Roitman (2015) and Agca and Igan (2013) who construct a series of the
consolidation measure for the years 2010 and 2011. The extension of the dataset is
based on the following three OECD reports: *Restoring Public Finances, 2011*, *Restor-
ing Public Finances, 2012 Update*, and *The State of Public Finances, 2015*. These
reports outline the economic situation, fiscal consolidation strategy and major consol-
idation measures for each of the OECD member countries. The country notes in each
report lay out each government’s rationale for pursuing fiscal adjustment and are used
to identify consolidation periods that were motivated by a desire for deficit reduction.

Table A3.2: Narrative Fiscal Shock, 2010-2014 (% GDP).

Country	2010	2011	2012	2013	2014
Australia	0.00	0.00	0.00	0.00	0.00
Canada	0.00	0.10	0.10	0.30	0.20
Germany	0.00	0.50	1.50	0.60	0.30
Denmark	0.00	0.90	0.00	1.10	-0.45
Spain	2.70	2.20	0.80	0.30	0.60
France	0.00	1.10	1.40	1.50	1.00
United Kingdom	0.60	1.20	1.00	1.00	-0.10
Italy	0.00	0.90	3.40	-0.49	0.47
Japan	0.00	0.00	0.00	0.00	0.00
Netherlands	0.00	0.30	0.70	2.10	1.90
Sweden	0.00	0.40	0.00	-0.60	-0.90
United States	0.00	0.00	0.00	1.00	0.00

Table A3.3: Controlling for CAPB (effect in year $t = 1$)

Specification	GDP		Consumption	
	High Debt	Low Debt	High Debt	Low Debt
Baseline ($\Delta CAPB_t$)	-1.13*** (0.50)	-0.21 (0.32)	-1.72*** (0.38)	-0.44 (0.46)
$CAPB_t$	-0.89** (0.52)	-0.64 (0.46)	-1.54*** (0.41)	-0.82 (0.74)

Notes: The table reports point estimates and robust standard errors clustered by country in parentheses. In each case the shocks are normalized so that the CAPB rises by 1% of GDP in year $t = 0$. *Significant at 16%; **significant at 10%; ***significant at 5%.

Table A3.4: Using BIS-Credit Data (effect in year $t = 1$)

Definition based on	GDP			Consumption		
	High Debt	Low Debt	Difference	High Debt	Low Debt	Difference
Baseline	-1.13*** (0.50)	-0.21 (0.32)	-0.92***	-1.72*** (0.38)	-0.44 (0.46)	-1.28***
BIS-credit data sample: 1980-2009	-0.85** (0.50)	0.01 (0.33)	-0.86***	-1.34*** (0.38)	-0.23 (0.48)	-1.11***

Notes: The table reports point estimates and robust standard errors clustered by country in parentheses. In each case the shocks are normalized so that the CAPB rises by 1% of GDP in year $t = 0$. *Significant at 16%; **significant at 10%; ***significant at 5%.

Table A3.5: Controlling for Linear Time Trend (effect in year $t = 1$)

Specification	GDP			Consumption		
	High Debt	Low Debt	Difference	High Debt	Low Debt	Difference
Baseline	-1.13*** (0.50)	-0.21 (0.32)	-0.92***	-1.72*** (0.38)	-0.44 (0.46)	-1.28***
Country-specific time trend	-0.90** (0.46)	-0.11 (0.33)	-0.79***	-1.52*** (0.42)	-0.24 (0.53)	-1.28***

Notes: The table reports point estimates and robust standard errors clustered by country in parentheses. In each case the shocks are normalized so that the CAPB rises by 1% of GDP in year $t = 0$. *Significant at 16%; **significant at 10%; ***significant at 5%.

Table A3.6: Leaving out Global Financial Crises (effect in year $t = 1$)

Specification	GDP			Consumption		
	High Debt	Low Debt	Difference	High Debt	Low Debt	Difference
Baseline	-1.13*** (0.50)	-0.21 (0.32)	-0.92***	-1.72*** (0.38)	-0.44 (0.46)	-1.28***
1978-2006 sample	-0.94*** (0.42)	-0.20 (0.31)	-0.74***	-1.58*** (0.38)	-0.39 (0.49)	-1.19***

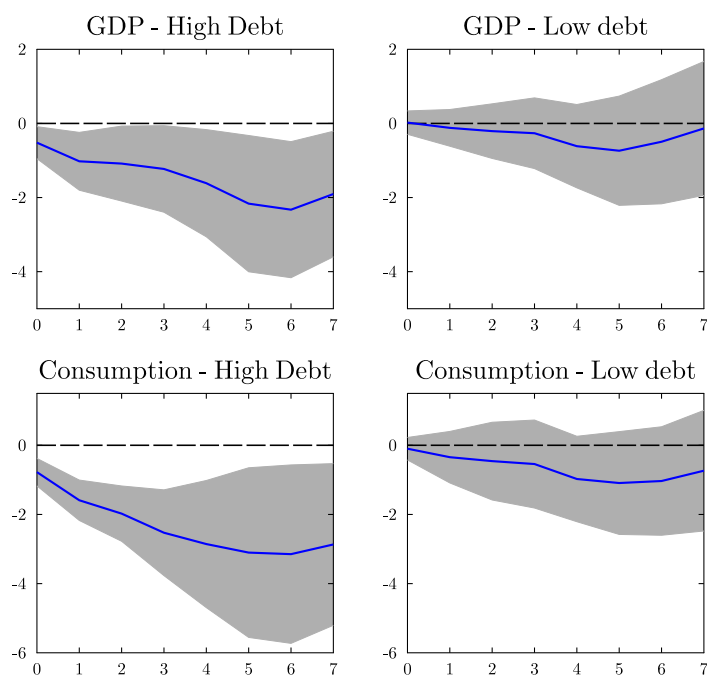
Notes: The table reports point estimates and robust standard errors clustered by country in parentheses. In each case the shocks are normalized so that the CAPB rises by 1% of GDP in year $t = 0$. *Significant at 16%; **significant at 10%; ***significant at 5%.

Table A3.7: Dropping one Country at a Time (effect in year $t = 1$)

Country excluded	GDP			Consumption		
	High Debt	Low Debt	Difference	High Debt	Low Debt	Difference
Baseline	-1.13*** (0.50)	-0.21 (0.32)	-0.92***	-1.72*** (0.38)	-0.44 (0.46)	-1.28***
Australia	-0.95** (0.53)	-0.25 (0.32)	-0.71**	-1.54*** (0.39)	-0.25 (0.32)	-1.10***
Canada	-0.93** (0.52)	-0.06 (0.28)	-0.87***	-1.61*** (0.44)	-0.29 (0.43)	-1.32***
Germany	-1.19*** (0.56)	-0.25 (0.33)	-0.94***	-1.78*** (0.43)	-0.44 (0.47)	-1.35***
Denmark	-1.09*** (0.51)	-0.36 (0.53)	-0.73**	-1.71*** (0.43)	-1.01* (0.65)	-0.70**
Spain	-1.04*** (0.46)	-0.07 (0.27)	-0.97***	-1.66*** (0.34)	-0.28 (0.39)	-1.38***
France	-1.16*** (0.55)	-0.22 (0.32)	-0.94***	-1.74*** (0.41)	-0.39 (0.43)	-1.35***
United Kingdom	-1.06** (0.55)	-0.27 (0.33)	-0.79**	-1.60*** (0.42)	-0.46 (0.48)	-1.14***
Italy	-1.89*** (0.62)	-0.29 (0.40)	-1.60***	-2.15*** (0.54)	-0.62 (0.58)	-1.53***
Japan	-0.89*** (0.36)	-0.29 (0.35)	-0.60***	-1.39*** (0.26)	-0.52 (0.51)	-0.87***
Netherlands	-0.88* (0.56)	-0.01 (0.26)	-0.87**	-1.55*** (0.44)	-0.15 (0.32)	-1.40***
Sweden	-1.22*** (0.59)	-0.36 (0.38)	-0.86**	-1.79*** (0.41)	-0.53 (0.55)	-1.26***
United States	-0.99** (0.51)	-0.17 (0.31)	-0.82**	-1.52*** (0.37)	-0.37 (0.47)	-1.15***

Notes: The table reports point estimates and robust standard errors clustered by country in parentheses. In each case the shocks are normalized so that the CAPB rises by 1% of GDP in year $t = 0$. *Significant at 16%; **significant at 10%; ***significant at 5%.

Figure A3.1: Estimating for Longer Horizon



Note: Cumulative changes (in per cent) in response to a shock of 1% of GDP to the cyclically-adjusted primary balance over $h = 0, 1, 2, 3, 4, 5, 6, 7$ years. The shaded areas indicate 90% confidence bands based on robust standard errors clustered by country.

4 Technology Shocks, Tax Cuts and their Impact on Private Household Debt

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ABSTRACT

This paper investigates the impact of total factor productivity shocks and tax innovations on household debt. Using vector autoregressions on US time series, we find that private borrowing increases substantially in response to both shocks. To account for these findings, we propose a DSGE model in which households' borrowing is limited by a collateral constraint, where durables play the role of collateral assets. By applying impulse response matching, we estimate structural parameters of the model and show that the model is capable of explaining the empirical observations.

Keywords: Financial Frictions, TFP shocks, Tax Shocks, Structural Estimation.

JEL Codes: E21, E32, E44.

4.1 Introduction

Since the beginning of the 1980s, total household debt increased substantially and almost doubled relative to GDP or private personal income in the US economy. This significant rise in household leveraging has led to a strand of literature studying the interaction between financial markets and the macroeconomy.²⁵ This study empirically investigates the impact of total factor productivity (TFP) shocks and tax innovations on household debt for the US economy and proposes a model with financial frictions that is capable of explaining the empirical observations.

²⁵Some influential studies are Kiyotaki and Moore (1997), Bernanke, Gertler, and Gilchrist (1998), and Iacoviello (2008).

It is widely agreed that introducing financial frictions into stochastic general equilibrium (DSGE) models changes the economic dynamics to shocks not just quantitatively but also qualitatively. Monacelli (2009) demonstrates that financial frictions are needed to account for the non-durable and durable consumption responses to a monetary policy shock as observed in the data. Based on vector autoregressions, Andrés, Boscá, and Ferri (2015) find that an expansionary government spending shock is followed by a significant and persistent increase in household debt. The authors propose a model in which private borrowing is limited to the value of the households' collateral in order to replicate the empirical impulse responses.²⁶ Based on these findings, this paper empirically shows that household debt moves procyclically in response to TFP shocks and tax innovations. Additionally, it is demonstrated that a DSGE model in which borrowing is limited by a collateral constraint can successfully account for these empirical results.

To study the impact on household debt, (i) the TFP series from Basu, Fernald, and Kimball (2006) and (ii) the Romer and Romer (2010) tax measure is incorporated into recursive SVARs. We select the TFP shock because technology improvements are one, among others, of the major drivers of the business cycle (e.g. Fisher (2006), Justiniano, Primiceri, and Tambalotti (2010)). Moreover, Mertens and Ravn (2012) empirically show that tax changes induce important impulses to US output fluctuations. Also, tax changes represent an important instrument for the fiscal authority to stimulate the economy. Thus, we study the effects of a fiscal and a non-fiscal shock.

So far, both of these shock series were mainly used to quantify their dynamic effects on variables like output, consumption or hours worked,²⁷ whereas this paper takes a closer look at how household debt evolves to changes in both measures.

Our empirical results suggest that increases in total factor productivity, as well as tax cuts, lead to a significant and persistent increase in household debt. Moreover, this paper finds that both shocks have expansionary effects on output, durable, and

²⁶In a similar vein, Eggertsson and Krugman (2012) theoretical show that the size of the government spending multiplier crucially depends on the degree of financial market imperfections.

²⁷For the TFP shock see, among others, Basu, Fernald, and Kimball (2006), Christiano, Eichenbaum, and Vigfusson (2004) and for the tax shock some prominent examples are Romer and Romer (2010), Mertens and Ravn (2012), Favero and Giavazzi (2012).

non-durable consumption. These findings indicate that the rise in economic activity in response to both shocks is partly financed by an increase in private borrowing. From a theoretical perspective, a positive debt response refutes consumption smoothing which assumes households to save in good times and borrow in bad. However, it is demonstrated that a theoretical model in which borrowing is limited by a collateral constraint as suggested by Monacelli (2009) produces such positive debt responses following both shocks. By applying impulse response matching it is then shown that this approach is capable of successfully explaining the empirical results.

Our proposed DSGE model is closely related to those used in the housing literature (Iacoviello, 2008) and in the literature on durable goods (Monacelli, 2009). The model economy is populated by two types of households, different in their willingness to postpone consumption into the future, which creates borrowers and lenders. Both agents earn after-tax labor income and receive utility from leisure and consuming a basket of durable and non-durable goods. The government purchases a stream of goods which is financed by distortionary labor income taxes and balances its budget every period by paying out lump-sum transfers. As the central building block of the model, borrowers face a collateral constraint so that the amount of newly issued private debt is restricted to a fraction of the value of their durable stock. Both economic shocks lead to an expansion in the modeled economy characterized by increases in output, non-durable consumption, and durable consumption. By assuming that the borrowing constraint holds with equality in the neighborhood of the steady-state, discount factors of the two types of households have to differ, as Iacoviello (2008) and Monacelli (2009) have shown.

To bring theoretical impulse responses as close as possible to the empirical data, deep model parameters are estimated. Instead of comparing the impulse responses from structural VARs to the theoretical responses from a model, this approach minimizes the distance between structural VAR responses run on the data and identical VAR responses run on simulated model data. Thus, the US data and the model simulations

are treated equally so that problems like small-sample biases or lag-truncation biases are avoided (Cogley and Nason, 1995; Kehoe, 2006).

Our results from the matching procedure suggest that the model can successfully account for the sizes and the hump-shaped patterns of the empirical dynamics in all four variables. In line with the empirical findings, the model produces persistent increases in household debt which last for more than 20 quarters. Moreover, the models' debt responses almost perfectly match the empirical counterparts. The point estimates of deep model parameters are in line with findings by previous studies (Mertens and Ravn, 2011). We estimate that almost 50% of all households are faced with a collateral constraint so that their ability to borrow to finance consumption is limited.

This study is a contribution to the existing literature in two dimensions. It is the first study giving a precise estimate for households' debt responses to technology improvements and tax cuts based on SVARs. Additionally, this paper contributes to the literature by showing that an estimated DSGE model with financial frictions matches the empirical responses of the major variable of interest, household debt, but also output, non-durable consumption, and durable consumption, to these shocks quantitatively.

The remainder of the paper is organized as follows. Section 2 presents the results from the SVAR estimation. Section 3 lays out the theoretical model. Section 4 describes the models' calibration and estimation strategy. Section 5 presents the results of the impulse response matching approach. Finally, the last Section 6 concludes.

4.2 Empirical Evidence

In this section, we present our data, estimation method, and SVAR results on the impact of technology shocks and tax cuts on total household debt and other main aggregates of interest.

4.2.1 Data and Identification

Our benchmark VAR consists of five variables. Apart from the main variable of interest, total household debt (d_t), we include output (y_t), non-durable consumption (c_t^n), consumption expenditures on durables (c_t^d), as well as one of the two shock measures, technology (z_t) or the tax innovation (τ_t). All variables are linearly detrended before estimation and enter the VAR in logs of real per capita, seasonally adjusted values.

Precise definitions and data sources are summarized in Table 4.1.

Table 4.1: Data Sources

Variable	Definition
y Output	Log of per capita Nominal gross domestic product divided by the GDP deflator
c^n Non-durable consumption	Log of per capita (personal consumption expenditures on non durables plus personal expenditures on services) divided by each individual price deflator
c^d Durable purchases	Log of per capita personal purchases of durable consumption goods divided by its deflator
d Total private debt	Log of per capita total private debt, divided by the consumer price index
τ Tax Shock	Romer and Romer (2010) exogenous tax shock
z Technology Shock	Basu, Fernald, and Kimball (2006) exogenous technology Shock

Notes: All data are linearly detrended and logs of real per capita, seasonally adjusted values and are obtained from FRED database. Full time series specific information and sources can be found in the appendix.

We measure the impact of technology shocks using the TFP series computed in Basu, Fernald, and Kimball (2006). This series is a Solow residual-based measure of technology corrected for labor and capital utilization, non-constant returns to scale, and imperfect competition.

To cope with the issue of endogenous and exogenous tax changes, we utilize the Romer and Romer (2010) tax measure.²⁸ The authors take a narrative approach to disentangling exogenous and endogenous tax change effects by analyzing presidential speeches, the Economic Reports of the President and reports of Congressional committees. Their resulting shock series is measured in changes in tax revenues relative to GDP, discounted to the day when the bill was signed, to avoid a misalignment of the data set and agents' economic choices, also called fiscal foresight.

Since the identification of the empirical model depends on the nature of the two shock series, i.e. if they are (strictly) exogenous, we perform Granger causality tests. In particular, we use these tests to find the suited VAR estimation method.

The results are summarized in Table 4.2. We find that the lags of output and durable consumption Granger cause TFP at the 95% significant level, which is in contrast to results reported in Basu, Fernald, and Kimball (2006)²⁹. When testing whether the Romer and Romer (2010) tax measure can be predicted by past observations of our main aggregates, Granger causality cannot be rejected either. Lagged values of output and durable consumption include information which help predict future tax changes.

Table 4.2: Granger Causality Test Results

	Basu, Fernald, and Kimball (2006) TFP series				Romer and Romer (2010) tax series			
	Sample	Obs	F-stat	p-value	Sample	Obs	F-stat	p-value
Output	1966–2014	191	2.395	0.042	1966–2007	164	2.487	0.025
Nondurables Cons.	1966–2014	191	1.723	0.146	1966–2007	164	1.486	0.186
Durables Purchases	1966–2014	191	3.673	0.006	1966–2007	164	2.871	0.011
Private Debt	1966–2014	191	0.967	0.426	1966–2007	164	1.121	0.303

Notes: Null hypothesis: The variable does not Granger cause the TFP / tax measure. Specification: six lags, linearly detrended data for output, non-durable consumption, durable consumption, hours worked, household debt (d) and TFP index as used in VARs. The Romer and Romer tax series is recoded as in Mertens and Ravn (2012).

²⁸Actual changes in tax rates or tax revenues are a linear combination of exogenous and endogenous tax changes, which would dilute the structural effect of tax innovations on total private debt. We hence seek to use a measure for exogenous tax changes only, rather than including automatic tax adjustments that co-move with the business cycle.

²⁹One explanation for the different results may be the different and shorter period considered by Basu, Fernald, and Kimball (2006).

Given these results, treating both measures as strictly exogenous series seems misleading and estimating exogenous VARs, as done in Basu, Fernald, and Kimball (2006) and Mertens and Ravn (2012), will not reveal the true impact of technology shocks and tax innovations on the variables of interest.

We acknowledge the fact of contemporaneous exogeneity of the two shock series by estimating VARs, in which the specific shock series are ordered first. This identification approach implies, that the TFP and the tax measure are contemporaneously unaffected by the other variables in the system while the subsequent variables have an impact through the lag structures. Our baseline SVAR takes the following form

$$X_t = A(L)X_{t-1} + u_t, \quad (4.1)$$

in which $X_t = [s_t, y_t, c_t^n, c_t^d, d_t]'$, $s_t \in \{z_t, \tau_t\}$. $A(L)$ is a lag polynomial of order 4 and the estimation includes a constant term. Finally, u_t denotes reduced form residuals, and their variance-covariance matrix is orthogonalized by Cholesky decomposition, and the VAR is estimated using ordinary least squares.

Both shocks enter our SVARs along with quarterly US data from 1966q1 to 2007q4 for the tax shock and 1966q1 to 2014q4 for the TFP shock. Due to data limitations of the tax measure the two samples have different lengths.

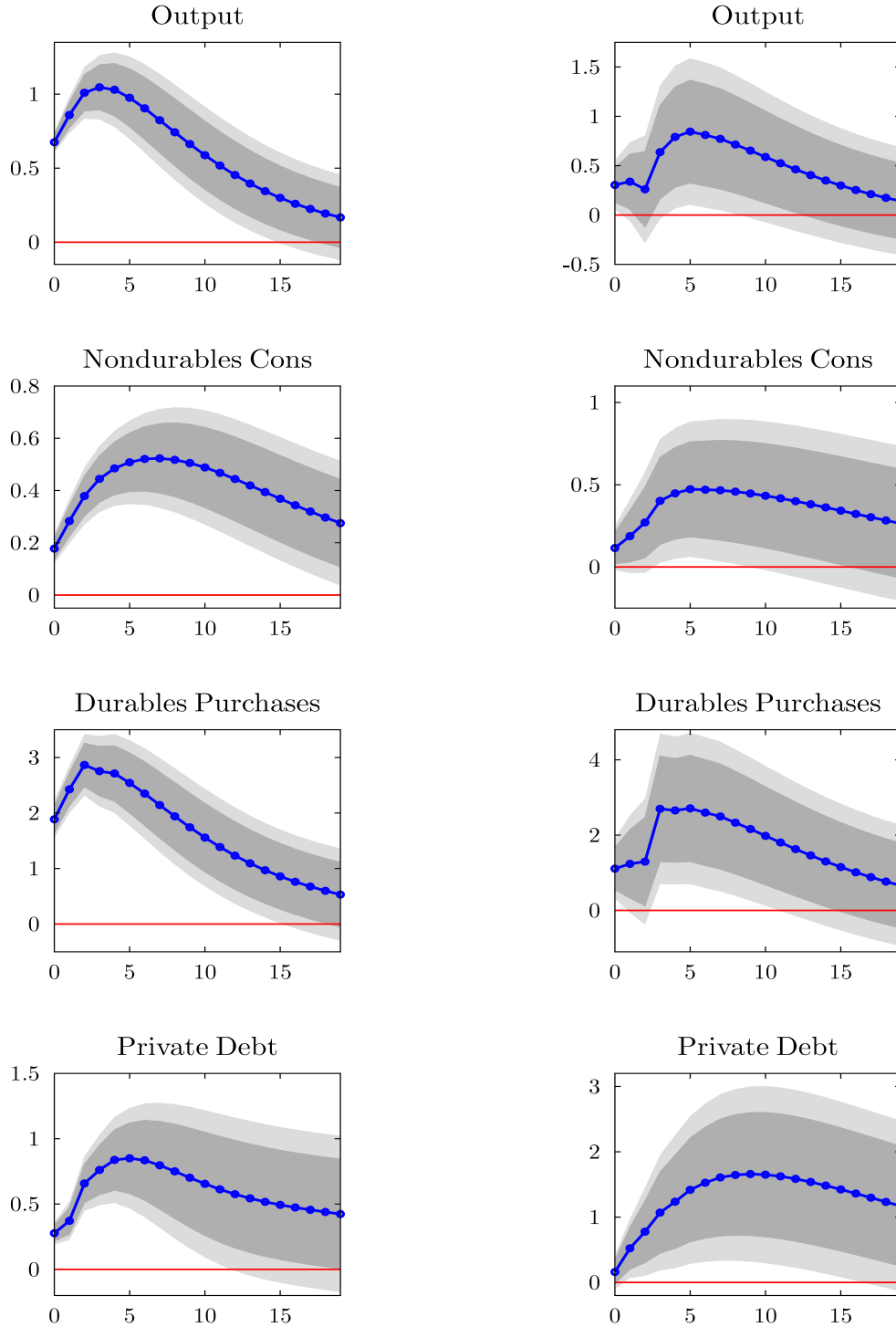
4.2.2 VAR Results

We produce one set of impulse responses for the TFP shock and one for the tax cut. The size of the technology shock is the equivalent to an increase of one percent in the Basu, Fernald, and Kimball (2006) measure for total factor productivity. The Romer and Romer (2010) structural tax shock is equivalent to a reduction of total tax revenues relative to GDP of one percentage point. We report impulse responses together with 68% (dark gray), and 84% (light gray) bootstrapped confidence bands, computed with 10.000 bootstrap replications. Figure 4.1 (a) depicts the results for the TFP shock and Figure 4.1 (b) those for the tax cut.

Figure 4.1: Impulse Responses SVAR Estimation

(a) TFP Shock

(b) Tax Shock



Notes: SVAR impulse responses to a one percent increase in TFP (a) and a one percentage point decrease in tax revenues over GDP along with 68% (dark gray) and 84% (light gray) bootstrapped centered confidence intervals with 10.000 bootstrapped replications. Reduced form residual variance-covariance matrices are Cholesky decomposed.

Both, the tax reduction and the increase in technological progress initiate an expansion characterized by hump-shaped dynamics in output, non-durable consumption, and durable expenditures. This boom is persistent, lasting for more than five years before the economy returns to its pre-shock level. While most of the variables do not change on impact when the economy faces a tax cut, the TFP shock influences the aggregates already on impact.

With respect to our primary variable of interest, household debt, we find that both shocks lead to a significant and persistent increase in private borrowing. This result indicates that the expansion in the economy is partly financed by a rise in household debt. For the TFP shock, household debt peaks after around 5 quarters, while for the tax cuts it converges back to pre-shock levels later. From a theoretical perspective, a sharp rise in household debt following both shocks is in contrast to consumption smoothing of households. This assumption would predict a fall in private borrowing in expansionary times as a buffer against future negative shocks.³⁰

Concerning the volatility of our endogenous variables, we can detect a clear pattern in responses to both exogenous innovations. Durable purchases and household debt react the strongest following both shocks.³¹ Non-durable consumption shows the smallest increases of all endogenous variables included in our VAR estimations.

These empirical findings are robust to alternative orderings of the variables in the VARs, less or additional lags, including hours in the estimation and, to the introduction of alternative variables. We find that the initially observed positive comovement between household debt and the other real variables remains intact.

4.3 Model

This section presents a DSGE model with financial frictions that we will use later to reproduce our empirical findings. The model consists of two types of households,

³⁰If we include the unemployment rate in our estimation, we find that unemployment is reduced after the two shocks. Thus, the rise in household debt is not caused by a decrease in household income.

³¹We interpret this strong comovement between durable purchases and household debt as justification for the borrowing constraint in our model which we describe in the next section.

a representative final goods firm, a monopolistically competitive intermediate goods sector, and a government sector.

4.3.1 Households

The model economy is populated by a continuum of infinitely-lived households that are heterogeneous in terms of their desire to save. Hence, a fraction χ of households becomes *lenders* (subscript l), while the remaining fraction $1 - \chi$ becomes *borrowers* (subscript b). Borrowing households face a collateral constraint which ensures that private borrowing is restricted to a certain amount of their stock of durables.

Lenders. Lending households' preferences are given by

$$E_0 \sum_{t=0}^{\infty} \beta_l^t \left(\frac{\Upsilon_{l,t}^{1-\sigma} - 1}{1-\sigma} - \gamma_l \frac{n_{l,t}^{1+\eta}}{1+\eta} \right), \quad (4.2)$$

where E_0 denotes the expectation operator conditional on all information available at time 0. $0 < \beta_l < 1$ is the lenders' specific discount factor, $\sigma > 0$ is a curvature parameter, $\gamma_l > 0$ is the preference weight that measures disutility of labor, $n_{l,t}$, and $\eta \geq 0$ is the inverse Frisch elasticity.

$\Upsilon_{l,t}$ denotes a consumption basket defined as

$$\Upsilon_{l,t} = c_{l,t}^\vartheta v_{l,t-1}^{1-\vartheta} - \psi_l c_{l,t-1}^\vartheta v_{l,t-2}^{1-\vartheta}, \quad (4.3)$$

where $c_{l,t}$ is consumption of non-durable goods and $v_{l,t-1}$ denotes the stock of durable goods held at the beginning of period t . $\vartheta \in [0, 1]$ measures the elasticity of substitution between non-durable and durable consumption, and $\psi_l \in [0, 1]$ governs the lenders' degree of habit persistence. We follow Mertens and Ravn (2011) by assuming that non-durable and durable consumption are complementary goods for households, which is assured by the specific functional form of (4.3).

Lending households maximize (4.2) with respect to their budget constraint given by

$$\begin{aligned}
c_{l,t} + x_{l,t} + b_{l,t} + d_{l,t} &\leq (1 - \tau_t)w_t n_{l,t} + (1 + r_{g,t-1})\frac{b_{l,t-1}}{\pi_t} \\
&+ (1 + r_{d,t-1})\frac{d_{l,t-1}}{\pi_t} + \frac{\Pi_t}{\chi} + tr_t,
\end{aligned} \tag{4.4}$$

where $\pi_t = p_t/p_{t-1}$ is the gross inflation rate and $b_{l,t}$ are the lender's holdings of one-period government bonds with interest $r_{g,t}$. Lenders receive after-tax labor income, $(1 - \tau_t)w_t n_{l,t}$, where τ_t is the labor income tax rate and w_t is the real wage rate which households take as given. $x_{l,t}$ represents purchases of new durable goods. In addition, lending households earn financial income, $(1 + r_{d,t-1})d_{l,t-1}$, from offering one-period private debt to borrowers at interest $r_{d,t-1}$ which is guaranteed to be repaid in the next period. tr_t denotes lump-sum transfers paid by the fiscal authority and Π_t/χ are the individual profits from owning intermediate goods firms.

The law of motion for the durable stock is given by

$$v_{l,t} = \left(1 - \frac{\phi_v}{2} \left(\frac{x_{l,t}}{x_{l,t-1}} - 1 \right)^2 \right) x_{l,t} + (1 - \delta_v)v_{l,t-1}, \tag{4.5}$$

in which δ_v denotes a constant depreciation rate and the parameter ϕ_v captures costs of adjusting the stock of durable goods. We choose this quadratic and convex functional form since it satisfies the properties generally imposed on adjustment costs (see, for example, Christiano, Eichenbaum, and Evans, 2005).³² Letting $\lambda_{l,t}$ be the lenders' Lagrange multiplier corresponding to their budget constraint, the first-order condi-

³²Let $\Phi(x_t/x_{t-1})$ be the general adjustment cost function. Then, convexity implies $\Phi(1) = \Phi'(1) = 0$ and $\Phi''(1) = \phi_v > 0$ which is assured by the functional form in (4.5).

tions (FOCs) for non-durable consumption, government bond holdings, hours worked, durable consumption, debt supply, and durable purchases are given by

$$c_{l,t} : \lambda_{l,t} = \vartheta \left(\Upsilon_{l,t}^{-\sigma} - \psi_l \beta_l E_t \Upsilon_{l,t+1}^{-\sigma} \right) \left(\frac{v_{l,t-1}}{c_{l,t}} \right)^{1-\vartheta}, \quad (4.6)$$

$$b_{l,t} : \lambda_{l,t} = \beta_l E_t \left\{ \lambda_{l,t+1} \frac{1+r_{g,t}}{\pi_{t+1}} \right\}, \quad (4.7)$$

$$n_{l,t} : \lambda_{l,t} (1-\tau_l) w_t = \gamma_l n_{l,t}^\eta, \quad (4.8)$$

$$v_{l,t} : \lambda_{l,t} q_{v,t} = \beta_l E_t \left\{ \lambda_{l,t+1} \left[\left(\frac{1-\vartheta}{\vartheta} \right) \frac{c_{l,t+1}}{v_{l,t}} + q_{v,t+1} (1-\delta_v) \right] \right\}, \quad (4.9)$$

$$d_{l,t} : \lambda_{l,t} = \beta_l E_t \left\{ \lambda_{l,t+1} \frac{1+r_{d,t}}{\pi_{t+1}} \right\}, \quad (4.10)$$

$$\begin{aligned} x_{l,t} : & 1 - q_{v,t} \left(1 - \frac{\phi_v}{2} \left(\frac{x_{l,t}}{x_{l,t-1}} - 1 \right)^2 - \phi_v \left(\frac{x_{l,t}}{x_{l,t-1}} - 1 \right) \frac{x_{l,t}}{x_{l,t-1}} \right) \\ & = \beta_l E_t \left\{ \frac{\lambda_{l,t+1}}{\lambda_{l,t}} q_{v,t+1} \phi_v \left(\frac{x_{l,t+1}}{x_{l,t}} - 1 \right) \left(\frac{x_{l,t+1}}{x_{l,t}} \right)^2 \right\}, \end{aligned} \quad (4.11)$$

where $q_{v,t}$ denotes the lenders' shadow value of new consumer durables. Equation (4.6) states that $\lambda_{l,t}$ equals the marginal utility of non-durable consumption. Equation (4.7) is the standard Euler equation for government bond holdings. Equation (4.8) sets the marginal rate of substitution between consumption and leisure equal to the after-tax real wage rate. Equation (4.9) shows that the shadow value of new consumer durables is equal to the expected discounted utility stream received from the durable stock (net of depreciation). Equation (4.10) sets $\lambda_{l,t}$ equal to the expected discounted utility stream of future debt interest rate payments. Equation (4.11) states that the change in consumer durables is related to the expected discounted value of current and future levels of $q_{v,t}$.

Borrowers. Preferences of borrowing households are defined as

$$E_0 \sum_{t=0}^{\infty} \beta_b^t \left(\frac{\Upsilon_{b,t}^{1-\sigma} - 1}{1-\sigma} - \gamma_b \frac{n_{b,t}^{1+\eta}}{1+\eta} \right), \quad (4.12)$$

in which $0 < \beta_b < 1$ is the specific discount factor of borrowers, $\gamma_b > 0$ is a scaling parameter measuring the borrowers disutility of labor, $n_{b,t}$. Again, $\Upsilon_{b,t}$ denotes a consumption basket defined as

$$\Upsilon_{b,t} = c_{b,t}^\vartheta v_{b,t-1}^{1-\vartheta} - \psi_b c_{b,t-1}^\vartheta v_{b,t-2}^{1-\vartheta}. \quad (4.13)$$

Here, $c_{b,t}$ denotes borrowers' consumption of non-durable goods and $v_{b,t-1}$ is the stock of durable goods held at the beginning of period t . $\psi_b \in [0, 1]$ measures the borrowers' degree of habit persistence.

The budget constraint of borrowing households is given by

$$c_{b,t} + x_{b,t} + (1 + r_{d,t-1}) \frac{d_{b,t-1}}{\pi_t} \leq (1 - \tau_t) w_t n_{b,t} + d_{b,t} + tr_t. \quad (4.14)$$

$x_{b,t}$ denotes borrowers' purchases of new consumer durables and $d_{b,t}$ is the amount of one-period-debt received from lenders which has to be repaid plus interest $r_{d,t-1}$ in the subsequent period. $(1 - \tau_t) w_t n_{b,t}$ denotes borrowers' after-tax labor income.

The borrowers' stock of durables accumulates according to

$$v_{b,t} = \left(1 - \frac{\phi_v}{2} \left(\frac{x_{b,t}}{x_{b,t-1}} - 1 \right)^2 \right) x_{b,t} + (1 - \delta_v) v_{b,t-1}. \quad (4.15)$$

As a central building block of our model, borrowing is endogenously determined by a collateral constraint, similar to the one used in Iacoviello (2008) and Monacelli (2009). The amount of debt that has to be repaid by borrowers in the following period, $d_{b,t}$, is the net-of-depreciation durable stock

$$d_{b,t} \leq \kappa(1 - \delta_v) v_{b,t}, \quad (4.16)$$

where $\kappa > 0$ denotes the share of borrowers' durable stock that can be used as collateral. This borrowing constraint implies two noteworthy points. First, by assuming that (4.16) holds with equality, β_b has to be smaller than β_l , and thus, borrowers hold a positive steady state amount of debt. Second, changes in the stock of durable goods

affect borrowing but also spending (of constrained households). The magnitude of this effect crucially depends on the size of κ .

The borrowers FOCs take the following expressions

$$c_{b,t} : \lambda_{l,t} = \vartheta \left(\Upsilon_{b,t}^{-\sigma} - \psi_b \beta_b E_t \Upsilon_{b,t+1}^{-\sigma} \right) \left(\frac{v_{b,t-1}}{c_{b,t}} \right)^{1-\vartheta}, \quad (4.17)$$

$$n_{b,t} : \lambda_{b,t} (1 - \tau_t) w_t = \gamma_b n_{b,t}^\eta, \quad (4.18)$$

$$v_{b,t} : \lambda_{b,t} q_{x,t} = \beta_b E_t \left\{ \lambda_{b,t+1} \left[\left(\frac{1 - \vartheta}{\vartheta} \right) \frac{c_{b,t+1}}{v_{b,t}} + q_{x,t+1} (1 - \delta_v) \right] \right\} + \mu_t (1 - \delta_v) \kappa, \quad (4.19)$$

$$d_{b,t} : \lambda_{b,t} = \beta_b E_t \left\{ \lambda_{b,t+1} \frac{1 + r_{d,t}}{\pi_{t+1}} \right\} + \mu_t, \quad (4.20)$$

$$\begin{aligned} x_{b,t} : 1 - q_{x,t} \left(1 - \frac{\phi_v}{2} \left(\frac{x_{b,t}}{x_{b,t-1}} - 1 \right)^2 - \phi_v \left(\frac{x_{b,t}}{x_{b,t-1}} - 1 \right) \frac{x_{b,t}}{x_{b,t-1}} \right) \\ = \beta_b E_t \left\{ \frac{\lambda_{b,t+1}}{\lambda_{b,t}} q_{x,t+1} \phi_v \left(\frac{x_{b,t+1}}{x_{b,t}} - 1 \right) \left(\frac{x_{b,t+1}}{x_{b,t}} \right)^2 \right\}, \end{aligned} \quad (4.21)$$

where $\lambda_{b,t}$ is the Lagrange multiplier on the borrowers budget constraint, μ_t denotes the Lagrange multiplier on the collateral constraint (4.16), and $q_{x,t}$ denotes the borrowers shadow value of new consumer durable purchases. Interpretations of equations (4.17), (4.18), and (4.21) are identical to those of lending households. The last term of (4.19) governs that the shadow value of new consumer durables is related to the marginal utility of relaxing the collateral constrained measured through the time-varying Lagrange multiplier μ_t . (4.20) shows that for positive values of μ_t the marginal utility of current consumption is larger than the marginal value of shifting one unit of consumption intertemporally. A higher value for μ_t induces a larger marginal benefit of increasing the stock of durable consumption goods which leads to a loosening of the collateral constraint to purchase additional current consumption.

4.3.2 Firms

The firm sector consists of a perfectly competitive final goods firm and a continuum of monopolistically competitive intermediate goods firms. Each intermediate goods firm $i \in [0, 1]$ produces a differentiated good $y_t(i)$ according to the production function

$$y_t(i) = z_t n_t(i), \quad (4.22)$$

where $n_t(i)$ denotes the quantity of labor services used by firm i . The technology level z_t is common across all operating firms and is assumed to follow an AR(1) process around its non-stochastic steady state value \bar{z} ,

$$\log(z_t) = (1 - \rho_z) \log(\bar{z}) + \rho_z \log(z_{t-1}) + \varepsilon_{z,t}, \quad (4.23)$$

in which $\varepsilon_{z,t}$ is *i.i.d* and $|\rho_z| < 1$. The representative final goods firm produces the final consumption good y_t , combining $y_t(i)$ units of each intermediate good, using the technology

$$y_t = \left(\int_0^1 y_t(i)^{\frac{\xi-1}{\xi}} di \right)^{\frac{\xi}{\xi-1}}, \quad (4.24)$$

where $\xi > 1$ is the elasticity of substitution between different intermediate goods. Profit maximization subject to (4.24) yields the demand function for intermediate good i ,

$$y_t(i) = y_t \left(\frac{p_t(i)}{p_t} \right)^{-\xi}, \quad (4.25)$$

where

$$p_t = \left(\int_0^1 p_t(i)^{1-\xi} di \right)^{\frac{1}{1-\xi}} \quad (4.26)$$

is the price index of the final good.

Each firm in the intermediate goods sector chooses its price level $p_t(i)$ to maximize the expected present value of real profits. Following Rotemberg (1982), each firm faces

quadratic adjustment costs which are assumed to take the functional form of Ireland (1997). Thus, real profits of firm i are given by

$$\Pi_t(i) = \left[\left(\frac{p_t(i)}{p_t} \right)^{1-\xi} - \frac{w_t}{z_t} \left(\frac{p_t(i)}{p_t} \right)^{-\xi} - \frac{\varphi}{2} \left(\frac{p_t(i)}{\bar{\pi} p_{t-1}(i)} - 1 \right)^2 \right] y_t, \quad (4.27)$$

where $\varphi > 0$ determines the adjustment costs and $\bar{\pi}$ is the steady state inflation rate.

Assuming symmetry in equilibrium, the optimality condition becomes

$$\varphi \left(\frac{\pi_t}{\bar{\pi}} - 1 \right) \frac{\pi_t}{\bar{\pi}} = (1 - \xi) + \xi \frac{w_t}{z_t} + E_t \left[\beta \frac{\lambda_{l,t+1}}{\lambda_{l,t}} \left(\frac{\pi_{t+1}}{\bar{\pi}} - 1 \right) \frac{\pi_{t+1}}{\bar{\pi}} \frac{y_{t+1}}{y_t} \right]. \quad (4.28)$$

In case of fully flexible prices, i.e. $\varphi = 0$, real marginal costs equal $(\xi - 1)/\xi$, which is the inverse of the firm's price markup.

4.3.3 Government

The government collects distortionary labor income taxes and issues new bonds to finance public spending, to service debt from last period, and to pay out lump-sum transfers to households. Hence, the government's budget constraint reads

$$g_t + tr_t + (1 + r_{g,t-1}) \frac{b_{t-1}}{\pi_t} = \tau_t w_t n_t + b_t, \quad (4.29)$$

where government spending g_t is a fixed fraction of aggregate output, and transfers tr_t adjust to balance the budget in every period. Following Mertens and Ravn (2011), we assume an AR(2) process for the tax rate around its non-stochastic steady-state value $\bar{\tau}$. The process is given by

$$\log(\tau_t) = (1 - \rho_{\tau,1} - \rho_{\tau,2}) \log(\bar{\tau}) + \rho_{\tau,1} \log(\tau_{t-1}) + \rho_{\tau,2} \log(\tau_{t-2}) - \varepsilon_{\tau,t}, \quad (4.30)$$

where $\varepsilon_{\tau,t}$ is *i.i.d.*, and $|\rho_{\tau,1} + \rho_{\tau,2}| < 1$.

Monetary policy is determined by a Taylor-type rule of the form

$$r_{g,t} = \bar{r}_g \left(\frac{\pi_t}{\pi^*} \right)^{\phi_\pi}, \quad (4.31)$$

where $\pi^* = \bar{\pi}$ is the inflation rate target and ϕ_π is the policy response to inflation deviations from its target.

4.3.4 Aggregation and Market Clearing

Aggregate variables are defined as the weighted average of the respective measures for each household type. Thus, we get

$$c_t = \chi c_{l,t} + (1 - \chi) c_{b,t}, \quad (4.32)$$

$$v_t = \chi v_{l,t} + (1 - \chi) v_{b,t}, \quad (4.33)$$

$$x_t = \chi x_{l,t} + (1 - \chi) x_{b,t}, \quad (4.34)$$

$$n_t = \chi n_{l,t} + (1 - \chi) n_{b,t}. \quad (4.35)$$

Credit and bond market clearing requires

$$\chi d_{l,t} = (1 - \chi) d_{b,t}, \quad (4.36)$$

$$b_t = \chi b_{l,t} \quad (4.37)$$

while the aggregate resource constraint is given by

$$c_t + x_t + g_t = \left[1 - \frac{\varphi}{2} \left(\frac{\pi_t}{\bar{\pi}} - 1 \right)^2 \right] y_t. \quad (4.38)$$

4.3.5 Equilibrium

A competitive equilibrium is given by the sequence of endogenous variables $\{y_t, c_t, c_{l,t}, c_{b,t}, v_t, v_{l,t}, v_{b,t}, x_t, x_{l,t}, x_{b,t}, \Upsilon_{l,t}, \Upsilon_{b,t}, n_t, n_{l,t}, n_{b,t}, d_{l,t}, d_{b,t}, b_t, b_{l,t}, \pi_t, tr_t, g_t, \lambda_{l,t}, \lambda_{b,t}, q_{v,t}, q_{x,t}, \mu_t, w_t, r_{d,t}, r_{g,t}\}_{t=0}^\infty$ that satisfy the households' first-order conditions, the firms' optimality conditions, the production function, the government budget constraint, the monetary policy rule, the stochastic processes, credit and bond market clearing, the aggregation identities, and the aggregate resource constraint, given the exogenous realizations of $\{z_t, \tau_t\}_{t=0}^\infty$.

To solve the model by a log-linear approximation around its deterministic steady state, we assume that all inequalities hold with equality in equilibrium.

4.4 Parametrization

To study whether our proposed model can account for the empirical findings, we estimate deep model parameters by applying an impulse-response matching approach as suggested by Cogley and Nason (1995). The set of parameters is partitioned into two subsets, $\Theta \equiv [\theta_1, \theta_2]$, where θ_2 contains the parameters to be estimated and θ_1 contains the parameters that are calibrated prior to estimation. The elements of the latter subset are fixed because they are either difficult to identify in model estimation procedures or are chosen to match certain steady-state targets observed in the data.

4.4.1 Calibration

One model period is set to be a quarter. We choose the lenders' discount factor to be 0.993, implying an annual steady-state interest rate of 3%, and follow Iacoviello and Neri (2010) by setting the borrowers' discount factor to 0.97 to induce a significant degree of impatience. The preference parameters determining disutility of work, γ_l and γ_b , are calibrated so that steady state hours worked equal 33% of individual time endowment. ϑ equals 0.75 which implies an aggregate steady state durable-to-non-durable-consumption ratio of 20%, in line with the corresponding number in the US during our sample period. The elasticity of substitution between intermediate goods, ξ , equals 11 implying a steady-state markup of $\xi/(\xi - 1) = 1.1$. For the debt-to-value ratio, κ , we again follow Iacoviello and Neri (2010) and choose 0.85 so that borrowing households can use 85% of their durable stock as collateral. Following Mertens and Ravn (2011), the depreciation rate of durable goods is set to 0.025 implying a steady state annual depreciation of 10%. We set the policy parameter in the Taylor rule, ϕ_π , to 1.5 as in Monacelli (2009). The steady state government-bonds-to-GDP, government-spending-to-GDP ratios, and the labor income tax rate equal 0.60, 0.18, and 0.28, respectively, as suggested by Trabandt and Uhlig (2011). Table 4.3 summa-

rizes the calibration of θ_1 .

Table 4.3: Model Calibration

Parameter	Value	Description	Target
β_l	0.993	Discount factor lenders	Ann. interest rate 3%
β_b	0.97	Discount factor borrowers	Small degree of impatience
γ_l		Preference parameter	SS hours of lenders to 0.33
γ_b		Preference parameter	SS hours of borrowers to 0.33
ξ	11.00	Elasticity of substitution	SS markup of 10%
δ_v	0.025	Depreciation rate durable goods	Mertens and Ravn (2011)
κ	0.85	Debt-to-value ratio	Andrés, Boscá, and Ferri (2013)
ϑ	0.75	Preference parameter	$\bar{X}/(\bar{C} + \bar{X}) = 0.20$
ϕ_π	1.50	Taylor rule parameter	Monacelli (2009)
\bar{B}/\bar{Y}	0.60	Government debt to GDP	Trabandt and Uhlig (2011)
\bar{G}/\bar{Y}	0.18	Government spending to GDP	Trabandt and Uhlig (2011)
$\bar{\tau}_n$	0.28	SS tax rate	Trabandt and Uhlig (2011)

4.4.2 Estimation

We estimate $\theta_2 = [\eta, \sigma, \psi_l, \psi_b, \phi_v, \chi, \rho_z, \varphi, \rho_{\tau,1}, \rho_{\tau,2}]$, by matching the impulse responses generated by the model to the empirical responses derived in Section 2. Estimating these parameters does not have an impact on the set of calibrated parameters in θ_1 . We target $\bar{n}_l = \bar{n}_b = 0.33$ so that γ_l and γ_b are endogenously determined.³³ We follow Cogley and Nason (1995) and Mertens and Ravn (2011), and treat model and data symmetrically. This implies that we use our model to simulate artificial samples and estimate impulse responses in exactly the same way as the empirical ones are obtained.³⁴

In particular, the model-generated impulse responses are constructed according to the following algorithm.

³³The two habit parameters have an impact on the steady-state value of \mathcal{T} and λ , but not on the first set of calibrated parameters θ_1 .

³⁴Using this approach avoids certain pitfalls of matching the theoretical impulse responses to the empirical ones, applied by e.g. Christiano, Eichenbaum, and Evans (2005) or Altig et al. (2011) as argued by Kehoe (2006) and Dupor and Liu (2003).

Algorithm 4.1 (Construction of model-generated IRFs) *For each of the two shocks, we take three steps:*

1. *Draw 100 sequences of innovations from the original shock series (with replacement) with a length of 168 periods for the tax shock and 195 for the TFP shock. Simulate the model for each draw so that there are 100 artificial samples. Each of these simulated datasets consists of the model counterparts to the SVAR time series.*
2. *Add a small (1e-6) white noise measurement error to each artificial time series to avoid stochastic singularity.*
3. *Estimate IRFs and take mean responses over all 100 replications for each artificial dataset by estimating (4.1).*

Let $\widehat{\Omega}_d$ be the vector of empirical moments and let $\widehat{\Omega}_m(\theta_2|\theta_1)$ be the vector of simulated moments estimated from the same SVAR as their empirical counterparts conditioned on θ_1 . Vector θ_2 then solves the following minimization problem,

$$\widehat{\theta}_2 = \arg \min_{\theta_2} \left[\left(\widehat{\Omega}_d - \widehat{\Omega}_m(\theta_2|\theta_1) \right)' W^{-1} \left(\widehat{\Omega}_d - \widehat{\Omega}_m(\theta_2|\theta_1) \right) \right], \quad (4.39)$$

where W is a positive-definite weighting matrix which we find by the following procedure. First, we approximate the covariance matrix of the empirical IRFs by bootstrapping. Instead of the full matrix, we only use its diagonal which displays the variances of the IRFs and set all off-diagonal elements to zero. Hence, we only put weight on moments that are observed in the data and force the estimation to exclude moments that are off-diagonal (see Cochrane, 2005, chap. 11). Finally, we use an estimate of the weighting matrix's asymptotic covariance matrix as proposed by Hall et al. (2012) to compute standard errors for θ_2 .

Table 4.4: Estimated Model Parameters

Parameter	Estimated Value	Standard Error	Description
σ	0.6664	0.1657	Utility curvature
η	0.2269	0.1127	Inverse Frisch elasticity
ψ_l	0.5672	0.3281	Habit parameter lenders
ψ_b	0.8394	0.0297	Habit parameter borrowers
ϕ_v	0.0921	0.0178	Durables adjustment cost
χ	0.5398	0.0425	Share of lending households
$\rho_{\tau,1}$	1.8611	0.0447	AR coefficient tax shock
$\rho_{\tau,2}$	-0.8745	0.0441	AR coefficient tax shock
ρ_z	0.9415	0.0165	AR coefficient tfp shock
φ	8.1276	0.7065	Rotemberg price adjustment

Notes: Standard errors are computed from an estimate of its asymptotic covariance matrix following Hall et al. (2012) and Mertens and Ravn (2011).

4.5 Results

Table 4.4 shows the parameter estimates of our model estimation. For the inverse Frisch elasticities, η , we estimate a value of 0.227. This value is somewhat lower than those typically assumed in the macroeconomic literature, whereas Iacoviello and Neri (2010) obtain a similar point estimate for a comparable model set-up. Our point estimate implies that labor supply of both agents reacts quite elastically to changes in the real wage rate. The degree of habit formation is larger for borrowing households than for lending ones, where the specific point estimates are in the range of values typically estimated in other studies (e.g. Christiano, Eichenbaum, and Evans, 2005).

The estimate of the durable adjustment cost parameter of 0.092 is lower compared to other studies (e.g. Mertens and Ravn, 2011). The estimated Rotemberg price adjustment coefficient, φ , takes a value of 8.128. The share of lending households in the economy is estimated to be 54%, consistent with estimates of the proportion of unconstrained consumers by Jappelli (1990), Kiley (2010).

Our estimates for the autoregressive shock parameters are 0.942 for the TFP shock, and 1.861, -0.875, for the tax shock. The degree of persistence of the tax process is similar to the one obtained by Mertens and Ravn (2011).

Figure 4.2 depicts the model dynamics to a one percent increase in total factor productivity (left panel) and to a one percentage point decrease in total tax revenues over output (right column) given the parameter estimates reported in Table 4.4 (dotted lines) along with the empirical estimates and its confidence bands from section 4.2.2.

As visible, the model can successfully account for the sizes and hump-shaped responses of the empirical counterparts. For almost all periods, the theoretical responses lie within the empirical confidence intervals. In line with the data, the strongest model responses can be observed for durable purchases and household debt, whereas non-durable consumption shows the smallest relative deviations following both innovations.

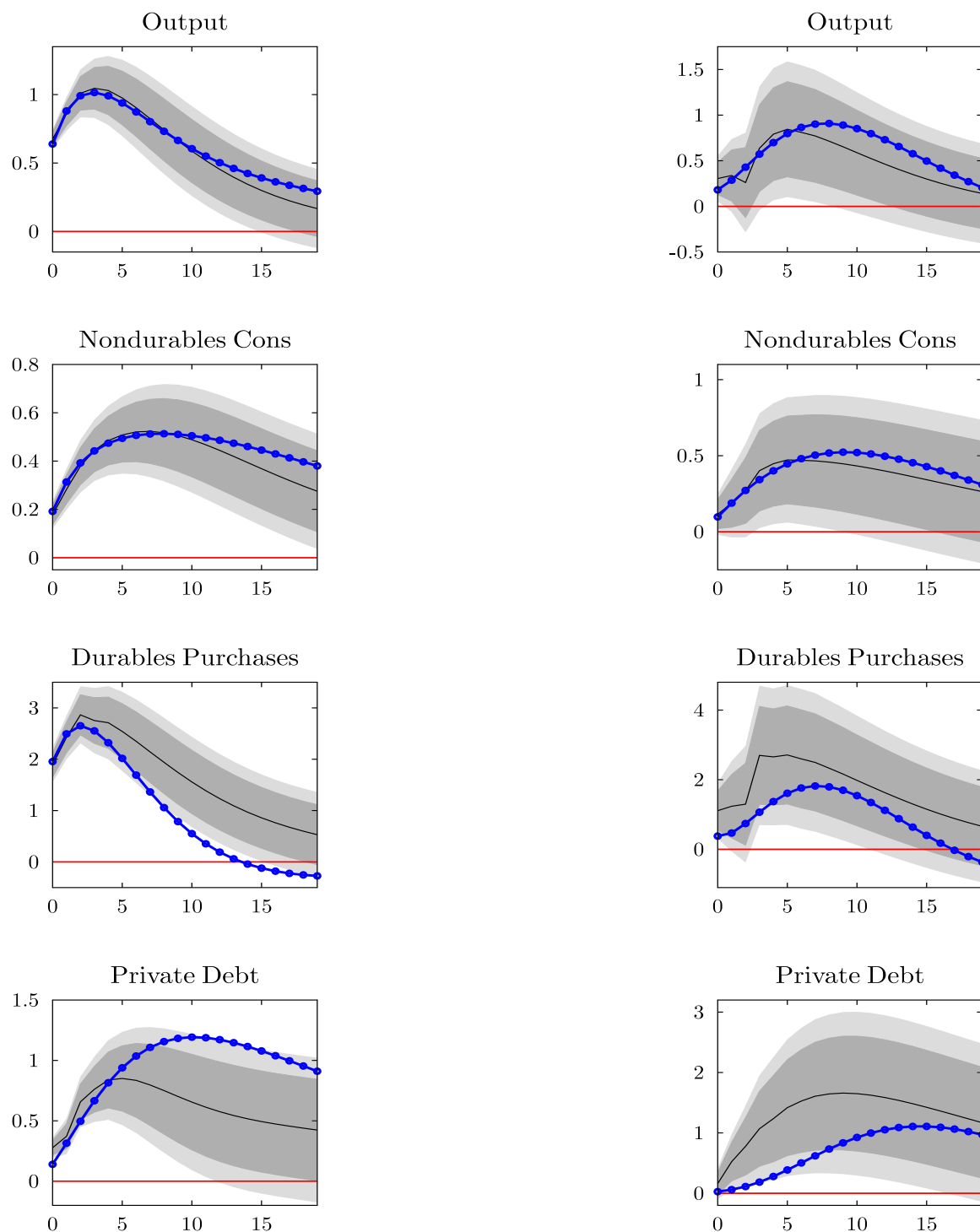
In panel (a) we observe that the model produces impact responses close to the empirical ones for the TFP innovation. The technology improvement leads to an expansion in the theoretical economy lasting for more than 20 quarters. The model's output, non-durable, and durable consumption responses reach its peaks slightly before the empirical counterparts. As durable consumption rises, the collateral constraint becomes less binding, such that borrowing households increase their private debt holdings. The model is able to capture the persistent increase in household debt as found in the data while the response lies within the 84% confidence bands for all periods. However, the model to some extent overestimates the debt response.

In line with our empirical findings, the model does not show any impact response for most of the variables after a tax reduction as can be observed in panel (b). The limited model responses can be explained by the estimated strong habits in consumption and positive durable adjustment cost which reduce the impact effects. For non-durable consumption, the model implied response reaches its peak after around seven periods, similar to the maximum empirical response. The specific maximum of the theoretical responses for output and durable consumption peak some quarters later than found for the empirical counterparts. The model underestimates the effect of tax cuts on durable purchases. Similar to the TFP shock, the model matches the households' empirical debt response quite well. The theoretical response falls within the empirical 68% confidence bands for most of the 20 periods. The increase in private borrowing following a tax

Figure 4.2: Empirical and matched Impulse Responses

(a) TFP shock

(b) Tax shock



Notes: This figure depicts VAR estimated impulse responses with actual data (solid line) along with 68% bootstrapped confidence bands (dark grey) and 84% confidence bands (light grey). The dotted lines denote matched impulse responses using our model.

reduction can be explained by the similar mechanism as described before for the TFP shock. The expansionary effects of the tax innovation lead to an increase in the stock of durables held by constrained agents such that, though the collateral constraint, private borrowing rises in response.

Our analysis suggests, that an estimated version of the model as described in section 4.3 is able to explain the empirical dynamics following technology improvements and tax reductions. When studying in more detail how private borrowing reacts, we find that the differences between theoretical and empirical responses are almost negligible.

4.6 Conclusion

The interrelation between financial market imperfections and macroeconomic outcomes is at the core of recent research. In this paper, we have studied the effects of TFP shocks and tax cuts on main aggregates for the US economy while taking a closer look at how households' borrowing decisions are affected by both innovations. We have selected these specific shocks because of their importance for business cycle fluctuations and as an important instrument for the fiscal authority to stimulate the economy.

By estimating SVARs, we find that both shocks lead to an expansion in the economy, characterized by significant increases in output, non-durable consumption, and durable consumption. Moreover, our results suggest that household debt rises substantially and in a hump-shaped manner in response to technology improvements and tax reductions.

In order to account for the empirically estimated comovement between economic activity and private borrowing, we have proposed a theoretically model with financial frictions similar to the one in Monacelli (2009). The model economy is populated by two household types, savers and borrowers, which differ in their willingness to postpone consumption into the future. Borrowers face a collateral constraint so that the amount of newly issued private debt is restricted to a fraction of their stock of durables.

We have estimated deep model parameters by matching the theoretically implied impulse response function to the empirical ones in response to both shocks. Our findings suggest that the proposed model successfully accounts for the sizes and hump-shaped

patterns of the empirical dynamics. With respect to our major variable of interest, household debt, the estimated model matches the empirical responses quite well. Estimated parameters are in line with findings in previous studies. Our estimates imply that almost 50% of private households do face a collateral constraint that restricts their optimal borrowing decision.

4.A Appendix

The data are obtained from FRED database and include the following data series.

Table A4.1: Full Data Sources

Variable	Definition	Code	Description
y	Output	GDP	Gross Domestic Product, Seasonally Adjusted
c^n	Non durable consumption	PCND	Personal Consumption Expenditures: nondurable goods, seasonally adjusted
		PCESV	Personal Consumption Expenditures: Services, Seasonally Adjusted
c^d	Durable purchases	PCDG	Personal Consumption Expenditures: Durable Goods, Seasonally Adjusted
h	Hours worked	h	Product of hours per worker and civilian non-farm employment divided by
d	total private debt	CMDEBT	Households and Nonprofit Organizations; Credit Market Instruments; Liability, Level, seasonally adjusted
N	Population	POP	Civilian Non institutional Population, Thousands of Persons
Price Deflators			
	GDP deflator	GDPDEF	Gross Domestic Product: Implicit Price Deflator
	Non durables deflator	DNDGRG3 Q086SBEA	Personal consumption expenditures: Nondurable goods
		DSERRG3 Q086SBEA	Personal consumption expenditures: Services
	Durables deflator	DDURRG3 Q086SBEA	Personal consumption expenditures: Durable goods
	CPI deflator	CPIAUCSL	Consumer Price Index for All Urban Consumers: All Items

Notes: This table gives FRED codes for the variables used in our estimation.

In addition to the FRED data series, we include the Romer and Romer (2010) tax series as τ , available at eml.berkeley.edu/~dromer/.

III Concluding Remarks

Concluding Remarks

This thesis has presented four essays that study the determinants of private household debt and the relation between private indebtedness and macroeconomic activity. Chapter 1 has shown that inequality and household debt are cointegrated of order one and therefore share a common trending relation. In Chapter 2, I have demonstrated that interpersonal comparison is an important driver of short-run credit movements. Chapter 3 has pointed out that the effects of fiscal consolidations crucially depend on the level of private indebtedness. In Chapter 4, I have presented a model with financial frictions that is able to replicate the empirical responses of household debt and other main macro aggregates to TFP shocks and income tax cuts.

Although my thesis contributes to the still growing literature on private household debt, there remain important questions which deserve further research. First, based on the findings of Chapter 3, it seems worth studying whether also the impact of fiscal expansions is significantly influenced by the level of private debt overhang. In future work, I want to elaborate on this topic and will try to contribute to the literature on state-dependent government spending multipliers (Auerbach and Gorodnichenko, 2012, 2013; Ramey and Zubairy, 2014). Second, the literature on the interrelation between monetary policy and household debt is still scanty. How monetary policy interventions affect households' debt positions and whether the effects of conventional or unconventional monetary policy are amplified by the level of private debt could be interesting questions for future research. Third, as shown by Mian and Sufi (2011, 2012) high private leverage ratios in the US are responsible for the long-lasting and persistent decrease in economic activity following the latest financial crisis. As a possible next research project, I would like to test whether private debt overhang also causes the slow recovery in the Euro Area.

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