# Saving Constraints, Debt, and the Credit Market Response to Fiscal Stimulus: Theory and Cross-Country Evidence

Jorge Miranda-Pinto*[1], Daniel Murphy†[2], Kieran James Walsh‡[3], and Eric R. Young§[4]

[1] University of Queensland
[2]University of Virginia Darden School of Business
[3]University of Virginia Darden School of Business
[4]University of Virginia and Zhejiang University

December 19, 2018

**Abstract**

We document that the interest rate response to fiscal stimulus (IRRF) is lower in countries with high inequality and high household debt. To interpret this evidence we develop a model in which households take on debt to maintain a minimum consumption threshold. Now debt-burdened, these households use additional income to deleverage. In economies with more debt-burdened households, increases in government spending tighten credit conditions less (relax credit conditions more), leading to smaller increases (larger declines) in the interest rate. To validate our mechanism we confirm that the consumption response to fiscal stimulus is lower in economies with high inequality or household debt. An implication of our theoretical and empirical results is that the sign of the debt-dependence of the effects of fiscal stimulus varies with credit conditions.

*Email: j.mirandapinto@uq.edu.au
†Email: murphyd@darden.virginia.edu
‡Email: walshk@darden.virginia.edu
§Email: ey2d@virginia.edu

# 1 Introduction

The size and length of the Great Recession renewed attention on fiscal policy as a stabilization tool. The design of optimal fiscal policy depends on an understanding of transmission mechanisms. The interest rate response to fiscal stimulus, which we call the IRRF, is of central importance, as it dictates the extent to which stimulus crowds out investment and (potentially) future output.

Despite the relevance of the interest rate channel, the literature has yet to offer clarity on how or why the interest rate responds to government spending or on what the IRRF can teach researchers about the underlying structure of the macroeconomy. This lack of attention and clarity may be due to an apparent conflict between theory and empirical findings. While standard theory (of both classical and Keynesian underpinnings) predicts that interest rates rise in response to government spending, studies based on the U.S. and U.K. tend to find a zero or negative effect on interest rates (e.g., Barro (1987) and, more recently, Ramey (2011) and Fisher and Peters (2010)). Related and also potentially puzzling is the evidence that government spending tends to be associated with local currency depreciation rather than appreciation (Ravn et al. (2012), Corsetti et al. (2012a), Faccini et al. (2016)).[1]

In this paper we use cross-country evidence, supplemented with U.S. regional micro data, to investigate the credit market effects of fiscal policy. We focus on government bond yields instead of short term interest rates to capture financial market conditions rather than the stance of monetary policy. We employ two approaches to identifying fiscal shocks. First, we follow Blanchard and Perotti (2002), who exploit relatively high frequency data and legislative lags to construct government spending innovations that are plausibly exogenous to current economic conditions. We also use the approach proposed by Auerbach and Gorodnichenko (2013), which, unlike that of Blanchard and Perotti (2002), takes into account the anticipation of government spending plans by using surveys of professional forecasters from OECD databases. We focus on the period before the Global Financial Crisis (GFC) since interest rates arguably respond to shocks differently in crisis periods.

We document that there is substantial heterogeneity in the IRRF across OECD countries, with approximately half of the countries experiencing a decline in government bond yields in response to an expansion of government consumption. Existing theory offers little guidance on the mechanisms that could account for these patterns. General equilibrium models

[1]The mechanism that would imply currency *appreciation* from government spending (vs. the depreciation seen in the data) is straightforward. Increased government spending crowds out private activity. The interest rate increases to clear the goods market, and higher rates attract foreign capital inflows, which appreciate the currency.

are generally unable to explain negative IRRFs (for longer term nominal government bond yields), and no theory of which we are aware has been proposed to account for heterogeneity in the IRRF (except with respect to the zero lower bound).

To shed light on the mechanisms responsible for this variation, we regress the IRRFs on country-level characteristics. We document that country-level income inequality and household debt are the strongest predictors of the IRRF. In particular, higher inequality and higher household debt are associated with a lower IRRF, both unconditionally and conditional on other potential country-level determinants of the IRRF. This result is surprising, given that one might expect high inequality or leverage to be associated with credit-constrained households with high marginal propensities to consume (see, for example, Huggett (1993), Aiyagari (1994), and Brinca et al. (2016)) that would, all else equal, push up the IRRF. The negative relationship between inequality or household debt and the IRRF suggests new theory is needed to understand the data.

To rationalize this evidence, we propose that high inequality or high household debt is often associated with low-income, debt-burdened households who have high marginal propensities to save (near-zero marginal propensities to consume (MPCs)) due to the desire to maintain a threshold level of consumption. If a household accumulates debt purely to stay above a minimum required consumption level, the household is then *debt-burdened* or *saving-constrained* in the sense that it uses additional income to delever rather than to further increase consumption. Our companion paper, Miranda-Pinto et al. (2018), lays out a theory of saving-constrained households and demonstrates that in a dynamic setting with incomplete markets, savings-constrained households exist in equilibrium (they do not fully precautionarily save to avoid the constraint in a calibrated model).[2] Here we demonstrate the macroeconomic implications of high-debt, saving-constrained households. The large debt is optimal conditional on the minimum consumption threshold, but it is too high relative to an environment without a minimum consumption threshold. The minimum consumption thresholds could represent, for example, aspects of current consumption (e.g., housing or auto maintenance) that are determined by prior decisions and costly to adjust in the short term, as in Chetty and Szeidl (2007), or unavoidable medical expenditures. Rather than move out of their homes or miss doctor appointments, households accumulate excessive debts that they are anxious to pay off.

We demonstrate the relevance of minimum consumption thresholds for credit markets in a two-period model with heterogeneous agents. The model illustrates in a simple setting how minimum consumption thresholds (saving constraints) generate an inverse relation-

[2]In Miranda-Pinto et al. (2018), we show that the existence of saving-constrained households is consistent with evidence from microdata. In particular, many high-debt/low-wealth households save all additional income (Sahm et al. (2015), Jappelli and Pistaferri (2014), Misra and Surico (2014)), consistent with the theory's predictions. We also demonstrate that our model can rationalize the joint dynamics of income and consumption in the Panel Study of Income Dynamics (PSID).

ship between inequality (and debt) and the IRRF. In our model a fraction of households are sufficiently poor that they hit the minimum consumption constraint in the first period. Government spending, along with progressive taxation, redistributes income to poor, saving-constrained households with low MPCs. More specifically, in producing government goods, the government hires and pays wages to workers, which are comprised of both high-debt (saving-constrained) low income agents (for whom the minimum threshold is binding) and unconstrained rich agents. Taxes are proportional to income, so wages associated with government production redistribute resources to the low-wealth households with zero MPCs. This redistribution to low-MPC households relaxes credit markets and puts downward pressure on the equilibrium interest rate, as government wages help poor workers delever. With higher inequality, more households are saving-constrained, household debt is higher, and government spending relaxes credits market more (tightens them less). This pattern offers an explanation for why the IRRF is lower in countries with higher inequality (and household debt).

To validate our mechanism we use cross-country and U.S. cross-county data to study how the private consumption response to government spending shocks depends on households' debt. Our mechanism implies that private consumption should increase less after fiscal shocks in countries or counties with higher household debt. The cross-country and cross-county evidence support this implication. We find that the four quarter cumulative response of consumption to government spending shocks is smaller in countries with high inequality or high household debt. With pre-Great Recession county level data for the U.S., we find that government spending increases auto registrations less in counties with high household debt.

Our empirical and theoretical results relate to a number of other strands of the literature. Recent empirical work documents determinants of fiscal output multipliers in cross-country settings (Brinca et al. (2016), Ilzetzki et al. (2013), Corsetti et al. (2012b)). We supplement this evidence by relating heterogeneity in the interest rate effect of fiscal policy to income inequality and household debt. Our empirical design is similar to Brinca et al. (2016), who find a positive relationship between fiscal multipliers and inequality. The authors explain their finding with a heterogeneous agent model with idiosyncratic risk and credit constraints. Unlike Brinca et al. (2016), who consider a broader set of countries including very unequal developing economies such as Brazil, Ecuador, and Colombia, we focus exclusively on OECD countries and restrict our analysis to the period before the GFC. To directly compare our results with Brinca et al. (2016), we analyze the relationship between our fiscal multiplier and inequality for the group of OECD countries. We find that the relationship is either non-existent or negative between inequality (or household debt) and fiscal multipliers. Moreover, our estimated negative relationship between the consumption response to fiscal shocks and inequality (or household debt) suggests that credit constraints are less relevant within

OECD countries during non-crisis times. Hence, comparing our results with those of Brinca et al. (2016), we conclude there is substantial heterogeneity between OECD and non-OECD economies and between crisis and non-crisis periods in the propagation of fiscal shocks.

Similarly, our evidence that the consumption response to government spending is *lower* in the presence of high household debt differs from recent evidence in Demyanyk et al. (2016) that consumer debt during the Great Recession was associated with higher consumption responses to fiscal stimulus. These findings can be reconciled by the fact that the Demyanyk et al. (2016) evidence is based on an episode in which credit conditions were incredibly tight, while our evidence is based on a longer span of time that exhibited looser credit conditions. To demonstrate the role of credit tightness in our theoretical framework, we introduce credit restrictions in our two-period model. When credit is sufficiently tight, poor households become credit-constrained rather than saving-constrained (they cannot even meet their minimum consumption threshold in the first period) and exhibit large MPCs. In that case, the consumption response to fiscal stimulus is increasing in inequality and debt, consistent with the evidence in Demyanyk et al. (2016) and with the theoretical predictions in Eggertsson and Krugman (2012). But under normal (looser) credit conditions, high-debt households are saving-constrained and exhibit low MPCs. We test this prediction using *pre-crisis* data across U.S. counties and find that consumption is indeed less response to fiscal stimulus in regions with more debt. This is consistent with the evidence in Demyanyk et al. (2016) that fiscal multipliers were, if anything, lower in high-debt cities in the mid-2000s. A key implication of our study is that not only is the effect of fiscal stimulus dependent on debt, but that the sign of this debt-dependence varies with credit conditions.

Finally, our evidence of negative IRRFs in a number of countries potentially helps resolve the puzzling finding of previous papers that expansionary government spending shocks are not clearly associated with exchange rate appreciations (see, for example, Corsetti et al. (2012a)). The standard Mundell-Fleming model predicts that exchange rates should increase as domestic interest rates rise, attracting capital inflows. Evidence against exchange rate appreciation has been interpreted as a rejection of Mundell-Fleming (Ravn et al. (2012)). Our paper offers a potential reconciliation between the data and the Mundell-Fleming interest-rate-channel of exchange rate movements.

The remainder of the paper proceeds as follows. Section 2 documents the relationship between the IRRF and inequality and household debt. Section 3 presents a qualitative theory of debt-burdened households to rationalize our findings. Section 4 presents several empirical validation exercises, including cross-county data for the United States. Section 5 concludes.

# 2 The interest rate response to fiscal stimulus

To estimate country-level fiscal shocks and the interest response to fiscal shocks (IRRFs), we collect quarterly data on real government consumption, real GDP, and interest rates across countries. Obtaining reliable country-level estimates of fiscal shocks requires a sufficient timespan of data. Therefore we limit our focus to OECD countries, most of which provide quarterly data that spans a period of over twenty years. The primary data source is the OECD. We supplement the OECD numbers with data from Haver when the Haver sample extends the OECD sample. A detailed description of the data used to estimate fiscal shocks is in Table 6 of our Appendix.

Our study focuses on government bond yields because they are the interest rate that is the most widely available for our sample. An advantage of examining yields on longer-dated bonds is that yields are not directly controlled by central banks but depend on credit conditions more generally. Our sample includes all OECD countries for which we observe government bond yields for at least 10 consecutive years prior to the end of our estimation period, 2007. Our baseline estimation period ends in 2007 in order to avoid structural breaks that may have been associated with the global financial crisis and to focus on the transmission mechanism of government spending shocks outside crisis times. In Appendix A we also examine data on shorter-term interest rates, which we refer to as policy rates. We use direct measures of central bank policy rates when available. For countries that do not have policy rate data, we use the short-term interest rate series in Ilzetzki et al. (2013). The policy rates for members of the European Monetary Union (EMU) are equal to European Central Bank rates.

## 2.1 Identifying Shocks to Government Consumption Expenditures

We identify government spending shocks following the approach in Blanchard and Perotti (2002). The key identification assumption is that, within a quarter, government spending is predetermined with respect to other macro variables. Hence government spending responds contemporaneously to its own shock but not to other shocks in the economy. Based on the delay in the political process that typically justifies this restriction, much of the literature has adopted the Blanchard-Perotti approach (e.g., Bachmann and Sims (2012), Auerbach and Gorodnichenko (2012), Rossi and Zubairy (2011), Brinca et al. (2016)).

Despite the widespread use of the Blanchard-Perotti approach and the plausibility of its identifying assumptions, there are potential limitations. If changes in government spending are anticipated, the Blanchard-Perotti approach will not capture the exogenous component of government spending (Ramey (2011)). To overcome this challenge, Ramey (2011) uses news about future defense spending to identify fiscal shocks. As Ilzetzki et al. (2013) point out, this approach is not viable when estimating fiscal shocks across countries. Data on military buildups on which the estimates are based are not available across countries, and even within the U.S. there is little variation in the military measure in the post-war period. Therefore, we adapt the Blanchard-Perotti approach. We acknowledge the potential limitations of this approach but note that the estimated effects of stimulus on interest rates are relatively consistent across empirical specifications, at least for the U.S. (see the discussion in Murphy and Walsh (2017)). As a robustness check, we also identify shocks using semi-annual data on forecast errors for government spending, as in Auerbach and Gorodnichenko (2013). We show in our Appendix that the main results of the paper also hold when we use the semi-annual government innovations from Auerbach and Gorodnichenko (2013).

We identify fiscal shocks independently for each country in our sample. To do so, we estimate

$$A_0 X_t = \sum_{j=1}^{4} A_j X_{t-j} + \varepsilon_t, \tag{1}$$

where $X_t = [G_t, Y_t, r_t]'$ consists of log real government final consumption expenditure $G_t$, log real GDP, and government bond yields $r_t$. $\varepsilon_t = \left[\nu_t, \varepsilon_t^2, \varepsilon_t^3\right]$ is a vector of structural shocks, and $v_t$ is the shock to government spending. The identifying assumption amounts to a zero restriction on the (1,2) and (1,3) elements of $A_0$. We use 4 lags of our endogenous variables. Unlike Blanchard and Perotti (2002), we do not have quarterly data on tax revenue for our sample.[3,4]

We estimate impulse responses of real GDP and interest rates to the fiscal shocks. For the purpose of our cross-country analysis, we summarize the information in the impulse responses by examining the cumulative 4-quarters impulse response to government consumption shocks. Let $\rho_h$ be the horizon $h$ impulse response of interest rates. To account for the variation across countries in the precision of the estimated impulse responses, we divide $\rho_h$ by the range between the upper (95%) and lower (5%) bounds of the bootstrap confidence intervals of the IRRFs, yielding $\hat{\rho}_h$.[5] Finally, the country-level interest rate response to a standard-deviation shock to government consumption is computed as:

$$IRRF = \sum_{h=0}^{3} \hat{\rho}_h.$$

[3]To explore how important is the omission of the tax revenue data, we check how the interest response to fiscal shocks in the VAR changes when tax revenue is included for the U.S. We find that the cumulative one year interest rate response is practically unchanged when tax revenue is added to the VAR. This is consistent with the findings in Ilzetzki et al. (2013) with respect to the output multiplier.

[4]We follow Auerbach and Gorodnichenko (2012) and estimate the VAR with the variables in log levels to preserve the cointegration relations. The fiscal shocks backed out from the VAR are stationary.

[5]The results are nearly identical without variance adjusting the IRRFs (see our Appendix). The variance adjustment helps limit the possibility is a small-sample setting that the results are driven by countries with large (in absolute value) but imprecisely estimated IRRFs. In a large-sample setting standard asymptotics would apply and the variance adjustment would be redundant.

Figure 1 depicts the substantial variation in the IRRF varies across countries. In half of the countries in the sample (14 countries), the response of interest rates to government consumption shocks is negative. In Switzerland a one standard deviation shock *increases* interest rates by 0.8 percentage points on average over four quarters. In the U.S., a standard deviation shock to government expenditure *decreases* interest rates by 0.32 percentage points.

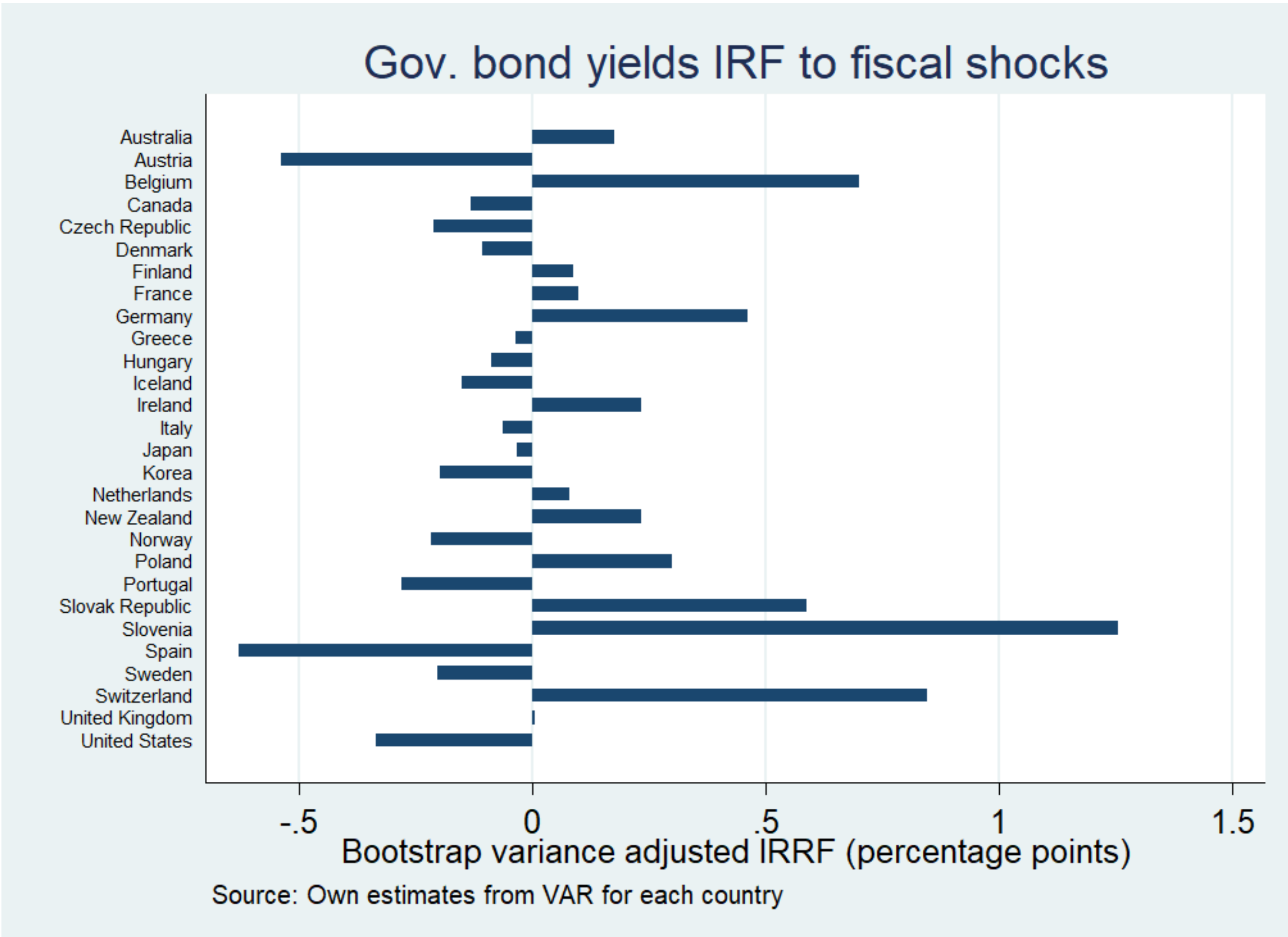


Figure 1
Effect of fiscal shocks on government bond yields across OECD Countries Blanchard and Perotti (2002)

Next we examine the country-level determinants of the IRRF. The exercise is similar in nature to Brinca et al. (2016), who demonstrate the fiscal multipliers are increasing in inequality in a sample that includes emerging economies. Based on their evidence, one might expect that interest rates are also increasing in inequality: if higher output multipliers are due to higher private spending propensities associated with poverty and inequality, then credit markets would be expected to tighten more in unequal countries. We find the opposite pattern in our sample of OECD countries. Inequality predicts a credit market loosening rather than tightening in response to government purchases. This pattern holds when conditioning on other potential determinants of the IRRF.

## 2.2 Determinants of the IRRF

Here we demonstrate that higher inequality and higher household debt are associated with a lower IRRF. Our measure of inequality is the ratio of the income of the richest 10 percent of the population to the income of the poorest 10 percent, which is provided by the OECD. For each country, we take the average over 2001-2013 (inequality is stable over time within countries). There is substantial cross-sectional dispersion in income inequality in our sample. The U.S. is the most unequal country of the sample with an average ratio of 6.2 , while Denmark has a ratio of 2.8. We also collect data on median household debt to assets (or income) from the OECD. These data are constructed from countries' microdata and in the case of debt-to-assets is only available for 23 countries. Figure 2 documents the unconditional relationship between the IRRF and inequality. We observe that the IRRF declines with inequality.

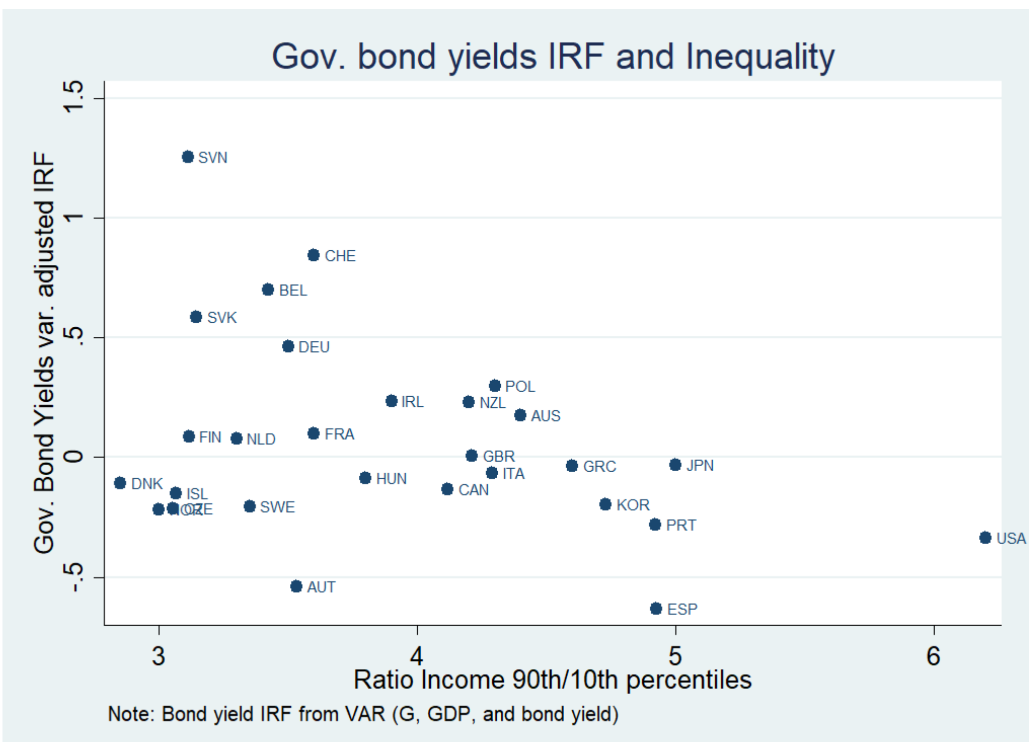


Figure 2
Gov. Bond Yield Response and Inequality

What else could account for this inverse relationship? One possibility is that monetary policy may be more accommodative of fiscal shocks in unequal countries. The same relationship does not hold when examining policy rates, suggesting that government spending relaxes credit markets relatively more in unequal countries, beyond any response of monetary policy to government spending shocks.[6] This is consistent with the evidence in Murphy and Walsh (2017) that monetary accommodation cannot fully account for the negative IRRF in the U.S.

To further isolate the role of inequality from central bank policy and other determinants, we regress the IRRF on measures of central bank independence and financial openness. We define a dummy variable for countries with inflation targeting scheme previous 2007 (see Carare and Stone (2003)). Our measure of financial openness, from Lane and Milesi-Ferretti (2007), is financial assets plus liabilities, over GDP. Mundell-Fleming predicts that countries that are more open to international financial markets have smaller or zero response of interest rates to fiscal shocks.

Motivated by Priftis and Zimic (2018) and Broner et al. (2018) we also control for the fraction of public foreign debt to GDP, obtained from the Quarterly Public Sector Debt statistics (IMF-World Bank). The authors show that fiscal multipliers are larger when government debt is externally financed due to a muted crowding-out of domestic credit markets. We calculate the average fraction of foreign public debt to GDP for the period 2002Q1-2017Q4. We only have this information for 19 of our 28 countries.

It is also possible that the relationship between inequality and the IRRF is due to a higher elasticity of output with respect to fiscal stimulus (perhaps due to slack, as in Murphy and Walsh (2017)) in countries with high inequality. In this case, interest rates would increase less in countries with more inequality, resulting in higher fiscal multipliers. One way to control for this channel is to include the fiscal multiplier in the regression. Here we define the fiscal multiplier as the cumulative multiplier over a horizon of one year, consistent with Ilzetzki et al. (2013):

$$\frac{\sum_{h=0}^{3} y_h}{\sum_{h=0}^{3} g_h}, \tag{2}$$

where $y_h$ and $g_h$ are the impulse responses of output and government consumption at horizon $h$.

Table 1 shows the dependence of the (variance adjusted) IRRF on inequality, conditional on these other determinants. We normalize our covariates, except inflation targeting, by their sample standard deviation.[7] We find that a one standard deviation increase in inequality is associated with a 15 basis point decline in the IRRF. The relationship is robust to controlling for countries' financial openness (column 2), inflation targeting (column 3), fiscal multiplier (column 4), and the fraction of government foreign debt to GDP (column 5). The same

[6] See Figure 9 in our Appendix.

[7] In non-reported results, we also control for countries' GDP per-capita, as in Brinca et al. (2016), to avoid that our results are driven by the degree of development of different countries. The results are unchanged

results hold if we instead use the point estimate – not variance adjusted – of the IRRF (see Table 7 in our Appendix).[8]

Table 1
Gov. Bond Yields Response and Country Characteristics

| VARIABLES | (1) IRRF | (2) IRRF | (3) IRRF | (4) IRRF | (5) IRRF |
|---|---|---|---|---|---|
| Income ratio 90th/10th | -0.145** | -0.133* | -0.151** | -0.132* | -0.187* |
| | (0.067) | (0.074) | (0.063) | (0.068) | (0.090) |
| Financial Openness | | 0.067 | | | |
| | | (0.069) | | | |
| Inflation Targeting | | | -0.283* | | |
| | | | (0.140) | | |
| Fiscal Mult. 4 qtrs | | | | 0.054 | |
| | | | | (0.068) | |
| Ext. G Debt to GDP | | | | | -0.029 |
| | | | | | (0.096) |
| Observations | 28 | 28 | 28 | 28 | 19 |
| R-squared | 0.125 | 0.150 | 0.246 | 0.141 | 0.195 |

* Note: This table presents the OLS coefficients of regressing the estimated (variance adjusted) IRRF – against income inequality (from OECD database), financial openness (from Lane and Milesi-Ferretti (2007)), inflation targeting dummy (from Carare and Stone (2003)), foreign government debt to GDP (from IMF-World Bank QPSD data), and the fiscal multiplier. The fiscal multiplier is estimated analogously to the IRRF. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

The theoretical model below offers an interpretation of the relationship between inequality and the IRRF. A key feature of the model is that inequality and household debt affect the IRRF through the same channels. Therefore here we also examine the relationship between the IRRF and the median household debt-to-assets ratio.[9] Table 2 shows that a one standard deviation increase in the household debt-to-assets ratio is associated with a 13 basis points reduction in the IRRF. This is also robust to adding controls (columns 2 to 4) and to alternative measures of the IRRF. In Tables 7-8 of our Appendix we show that the same results hold for the non-variance adjusted IRRF.[10]

[8]The coefficient on inflation targeting is negatively and statistically significant, as expected. The coefficient for foreign public debt to GDP is qualitatively consistent with the predictions in Priftis and Zimic (2018) and Broner et al. (2018), however, it is not statistically different from zero.

[9]Similar results, although less precise hold when we use household debt to income ratio. The correlation between household debt-to-assets and household debt-to-income is 0.75.

[10]In Table 2 we do not add the fraction of foreign public debt because our sample size is substantially reduced to only 15 countries. In addition, from Table 1 we observe that, while this control variable has the expected negative sign, it is not particularly relevant in accounting for the cross-country heterogeneity in the IRRF.

Table 2
Gov. Bond Yields Response and Country Characteristics

| VARIABLES | (1) IRRF | (2) IRRF | (3) IRRF | (4) IRRF |
|---|---|---|---|---|
| HH debt to assets | -0.132* | -0.141* | -0.130* | -0.155* |
| | (0.070) | (0.072) | (0.069) | (0.074) |
| Financial Openness | | 0.063 | | |
| | | (0.046) | | |
| Inflation Targeting | | | -0.198 | |
| | | | (0.156) | |
| Fiscal Mult. 4 qtrs | | | | 0.086 |
| | | | | (0.081) |
| Observations | 23 | 23 | 23 | 23 |
| R-squared | 0.102 | 0.126 | 0.161 | 0.135 |

* Note: This table presents the OLS coefficients of regressing the estimated four quarters variance adjusted IRRF – against median household debt to assets ratio (from OECD database), financial openness (from Lane and Milesi-Ferretti (2007)), inflation targeting dummy (from Carare and Stone (2003)), and the fiscal multiplier. The fiscal multiplier is estimated analogously to the IRRF. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

To summarize our results, the interest rate response to government purchases is heterogeneous across countries and is inversely related to inequality and household leverage. Below we propose a model in which high inequality and high debt are associated with a large fraction of low-income households with high propensities to save (low MPCs). Government consumption redistributes resources to these low-income households and relaxes credit markets.

# 3 Theory: Saving-constrained households, debt, and interest rates

In a standard economy with a representative agent or aggregation, the interest rate is invariant to the income/debt distribution, and the interest rate unambiguously rises in response to government spending shocks: when the government uses resources today, market clearing requires the interest rate to rise to induce the representative agent to forgo consumption today for consumption tomorrow. Therefore, market frictions are required for interest rate responses that are potentially negative and depend on the income distribution.

Here we develop a framework in which the distribution of debt is crucially important for the transmission of fiscal policy. The simplest version of our model predicts that the fraction of households with low MPCs (high marginal propensities to save) is increasing in the amount of credit/debt in the economy. When there are sufficiently high levels of debt outstanding, there is a strong desire to delever out of additional income, and increases in income from the government pushes down the equilibrium interest rate.

We are of course not the first to explore the importance of debt for the transmission of fiscal policy (e.g., Eggertsson and Krugman (2012)). What is unique about our framework is that indebted households are not credit constrained but instead want to delever out of additional income (they are "savings constrained"). How does such a situation emerge?

We consider a friction – a minimum consumption level – that results in some agents having high debt and unable to lower consumption to delever. A minimum consumption level could represent, for example, a subsistence level of consumption or aspects of current consumption (e.g., rent) that are determined by prior decisions and costly to adjust in the short-term (similar to "consumption commitments" in Chetty and Szeidl (2007)). It also represents unanticipated events such as medical emergencies and auto repairs that require high expenditure (or a substantial utility cost), as discussed in Miranda-Pinto et al. (2018). When low-income households experience an adverse shock, such as an increase in the minimum expenditure threshold or a decline in income, they borrow more than they would in the absence of the minimum consumption constraint. We refer to this excess debt as debt burden: households pay off the debt as soon as they receive additional income. In this sense, these households are *saving*-constrained (rather than credit-constrained).[11]

## 3.1 Model

Suppose there are two agent types, rich$(r)$ and non-rich $(p)$. The measure of non-rich agents is $\pi \in [1/2, 1)$, and the measure of rich agents is $1-\pi$. Each agent elastically supplies up to $\overline{L}$ units of labor in each period, of which there are two: $t \in \{0, 1\}$.

In each period, there is a representative private firm that solves

$$\Pi = \max_{\ell} \left( A\ell^{\alpha} - w\ell \right),$$

where $w$ is the wage, which is stuck, and $0 < \alpha < 1$. Given $w$, firm labor demand is $\ell^* = (w/(\alpha A))^{1/(\alpha-1)}$. We assume that (1) $\overline{L} > \ell^*$, (2) the firm randomly hires amongst the agents, and (3) $A = (w/\alpha)^{\alpha}$ (a simplifying normalization). Therefore, firm and worker optimization imply that $\Pi + w\ell^* = A\ell^{*\alpha} = 1$, that $\ell^* = \alpha/w$, and that each agent's private sector labor income is $w\ell^* = \alpha$, a fraction $\pi$ of which goes to non-rich agents. Moreover, since $\ell^* < \overline{L}$

[11] In Miranda-Pinto et al. (2018), we introduce the theory of saving-constrained households in a dynamic incomplete-market heterogeneous agent setting and demonstrate that the theory can rationalize otherwise unexplained features of microdata. Here we embed saving constraints in a general equilibrium framework that allows us to explore the interrelationships between debt, inequality, and the effects of fiscal policy on credit markets.

there is slack in the labor market in the sense that each agent is willing to supply more labor than the private sector is willing to hire at the stuck wage $w$.

In $t=0$, the government also hires the agents (again, randomly across types). Specifically, the government demands $\widetilde{G}=G/w<\overline{L}-\ell^*$ units of labor, which the agents are willing to supply since $\widetilde{G}+\ell^*<\overline{L}$. The government uses the workers to produce government goods and effectively buys these goods from itself. For the purposes of national accounting, these public purchases are valued at their cost. So, $G=\widetilde{G}w=\pi\widetilde{G}w+(1-\pi)\widetilde{G}w$ is both the public wage paid to each agent and the value of government purchases in the national accounts. GDP or national income is, in the two periods,

$$
\begin{aligned}
Y_0 &= \Pi+w\ell^*+w\widetilde{G}=A\ell^{*\alpha}+G=1+G \\
Y_1 &= \Pi+w\ell^*=A\ell^{*\alpha}=1
\end{aligned}
$$

We assume that the rich collectively own half of firm profits. Thus, the total private sector pre-tax income of the rich is $\Pi/2+(1-\pi)w\ell^*$, while the income of a rich individual is $y^r=\Pi/(2(1-\pi))+w\ell^*$. Similarly, the private sector pre-tax income of a non-rich individual is $y^p=\Pi/(2\pi)+w\ell^*$, so $(1-\pi)y^r+\pi y^p=1$. A useful feature of this setup is that a single parameter, $\pi$, governs inequality. As $\pi$ varies between 1/2 and 1, total private income is fixed at $\Pi+w\ell^*=1$. However, since the poorest 50% of agents are always non-rich, the total private pre-tax income of the richest 50% of agents is

$$
\Pi+w\ell^*-\frac{1}{2}\left(\frac{\Pi}{2\pi}+w\ell^*\right),
$$

which is monotonically increasing in $\pi$. Also, as $\pi\to 1$, half of firm profits are owned by an increasingly small fraction of agents.

In the first period, the agents and the government trade zero net supply bonds at gross interest rate $R$. The government pays for purchases with a flat proportional tax $\tau$ on private income in the second period. Since $(1-\pi)y^r+\pi y^p=1$, the government budget constraint is

$$
RG=\tau.
$$

The problem of an arbitrary agent of type $i\in\{r,p\}$ is

$$
\begin{aligned}
&\max_{c_0,c_1}\{\log(c_0)+\log(c_1)\} \text{ subject to} \\
&(i): c_0+\frac{1}{R}c_1=y^i+\frac{1}{R}y^i(1-\tau)+G \\
&(ii): c_0\geq \underline{c},
\end{aligned} \tag{3}
$$

where $\underline{c}$ is the minimum consumption level. Recall that $G=\widetilde{G}w$ is wage income from government work, and $y^i$ includes both private profits and wages. Since taxes are proportional to private income but government wages are uniform across agents, fiscal policy redistributes from rich to non-rich.

Under the above assumptions, equilibrium with slack in the labor market consists of an interest rate $R$, agent consumption, and taxes $\tau$ such that goods markets clear ($\pi\left(c_0^p, c_1^p\right) + (1-\pi)\left(c_0^r, c_1^r\right) = (1,1)$), consumption solves the agents' problems (3) given prices and taxes, and the government budget constraint is satisfied ($RG = \tau$).[12]

To restrict attention to the case of interest in which $c_0^r > c_0^p = \underline{c}$, we impose the following additional parameter restriction:

$$\frac{\Pi}{4}\left(\frac{2\pi-1}{\pi}\right)G + \frac{\Pi}{2\pi} + w\ell^* \leq \underline{c} < 1.$$

Since $\Pi/(2\pi) + w\ell^* < 1$ for $\pi > 1/2$, there exists $\underline{c} \in (0,1)$ satisfying this condition provided $\pi > 1/2$ and $G$ is sufficiently small. In this case, optimal rich consumption satisfies

$$c_0^r = \frac{1}{2}G + \frac{1}{2}y^r\left(1 + \frac{1}{R}(1-\tau)\right),$$

which after plugging in the government budget constraint becomes

$$c_0^r = \frac{1}{2}(1-y^r)G + \frac{1}{2}y^r\left(1+\frac{1}{R}\right). \tag{4}$$

Finally, imposing market clearing ($\pi c_0^p + (1-\pi)c_0^r = 1$) and $y^r = \Pi/(2(1-\pi)) + w\ell^*$, we get

$$\begin{aligned}\frac{1}{R} &= \frac{2(1-\pi\underline{c})}{\frac{\Pi}{2} + w\ell^*(1-\pi)} - \frac{1-\left(\frac{\Pi}{2(1-\pi)} + w\ell^*\right)}{\frac{\Pi}{2(1-\pi)} + w\ell^*}G - 1 \\ &= \frac{2(1-\pi\underline{c})}{(1-\pi)y^r} - \frac{1-y^r}{y^r}G - 1\end{aligned}$$

It immediately follows that

$$\frac{\partial^2(1/R)}{\partial G\partial\pi} > 0,$$

implying

**Proposition 1** $\frac{\partial^2 R}{\partial G \partial \pi} < 0$

In particular, the impact of $G$ on $R$ is declining in inequality. Government spending redistributes from high MPC to low MPC households, which relaxes credit markets more when the economy is populated by a larger fraction of debt-burdened households. Note, however, that in this stripped-down model increasing government purchases actually unambiguously decreases the interest rate, contrary to standard intuition. This is because here government

[12] The government goods market clears for free since, by assumption, the government consumes whatever it produces. The labor market doesn't clear since each agent is willing to supply $\overline{L}$, while at stuck wage $w$ private and public firms only demand $\ell^* + \widetilde{G} < \overline{L}$ units of labor from each agent.

spending destroys no resources. However, it is trivial to include government waste by assuming that government consumption/production $G$ requires an input $\omega G$ of the consumption good, meaning the public budget constraint becomes $G(1+\omega)R = \tau$. In that case, the sign of $\partial R/\partial G$ is positive when $\omega$ is high, but $\partial^2 R/(\partial G \partial \pi) < 0$ still holds.[13]

Therefore, a theory with saving constraints suggests that high inequality is associated with a weaker or even negative response of interest rates to government spending. The same is true with respect to debt: at $t = 0$ a non-rich agent is borrowing $\underline{c} - (y^p + G)$, which is increasing in $\pi$, when $\underline{c} = c_0^p$. This immediately implies that total private debt, $\pi(\underline{c} - (y^p + G))$, is also associated with inequality and a low IRRF.

The credit market relaxation in response to government purchases occurs entirely through the interest rate. Since private output is fixed (and under the assumption that the government does not purchase private-sector output, so is aggregate consumption), there is no quantity adjustment from credit market relaxation. In a more complicated setup with elastic private-sector output, however, the adjustment will occur through both prices (the interest rate) and quantities (consumption). In our setting, for example, aggregate desired consumption is $\pi c_0^p + (1-\pi)\, c_0^r$. By equation 4, it follows that aggregate desired consumption (prior to imposing market clearing) is

$$C(R) = \pi \underline{c} + (1-\pi)\left[\frac{1}{2}(1-y^r)\, G + \frac{1}{2} y^r \left(1 + \frac{1}{R}\right)\right],$$

and hence, since $y^r = \Pi / (2(1-\pi)) + w\ell^*$

$$\frac{\partial^2 C}{\partial G \partial \pi} < 0.$$

Therefore, an implication of the theory with saving-constrained households is that the relationship between inequality (and debt) and the consumption response to fiscal stimulus (CIRF) is non-positive. In our simple theoretical setting there is no relationship due to simplifying assumptions about the supply side of the economy, but is a setting with elastic private-sector output, the predicted relationship will be negative. Below we confirm that the CIRF is, if anything, inversely related to inequality and debt.

[13] In a neoclassical version of the model with fully flexible prices and government spending that crowds out private consumption, the IRRF is declining in inequality but the interest rate always rises in response to government purchases

# 4 Testing an implication of the model: The consumption response to fiscal stimulus

In this section, we examine how the consumption response to fiscal shocks depends on household debt. A useful feature of consumption (relative to government bond yields) is that it varies across regions within a country. Therefore, in additional to examining cross-country evidence as in Section 2, we also examine U.S. cross-county evidence by exploiting regional variation in household debt and consumption.

## 4.1 Cross-Country consumption response and debt

Here we test the theory's prediction that the relationship between the manifestations of savings constraints (inequality and debt) and the consumption response to fiscal stimulus is non-positive. Similar to section 2, we identify fiscal shocks independently for each country in our sample. To do so, we estimate equation 1 using $X_t = [G_t, Y_t, C_t]'$. $X_t$ consists of log real government spending $G_t$, log real GDP, and log real private consumption $C_t$. $\varepsilon_t = \left[\nu_t, \varepsilon_t^2, \varepsilon_t^3\right]$ is a vector of structural shocks, and $v_t$ is the shock to government spending. The identifying assumption follows Blanchard and Perotti (2002) as in section 2.1.

We also summarize the information in the impulse responses by examining the cumulative 4-quarters impulse response to government consumption shocks. Let $\rho_h^c$ be the horizon $h$ impulse response of consumption. To account for the variation across countries in the precision of the estimated impulse responses, we divide $\rho_h^c$ by the range between the upper (95%) and lower (5%) bounds of the bootstrap confidence intervals of the IRRFs, yielding $\hat{\rho}_h^c$. Finally, the country-level consumption response to a standard-deviation shock to government consumption is computed as:

$$CIRF = \sum_{h=0}^{3} \hat{\rho}_h^c.$$

The pattern in Figure 3 is consistent with credit market relaxation in response to government purchases. There is a negative relationship between inequality (and household debt) and the cumulative four quarters response of private consumption to government spending shocks. Table 3 shows that the relationship between CIRF and household debt to income is negative and statistically significant, regardless whether we adjust the CIRF by its variance or whether we include other controls in the regression.[14]

[14]Similar, but less precise estimates hold when we use inequality rather than household debt.

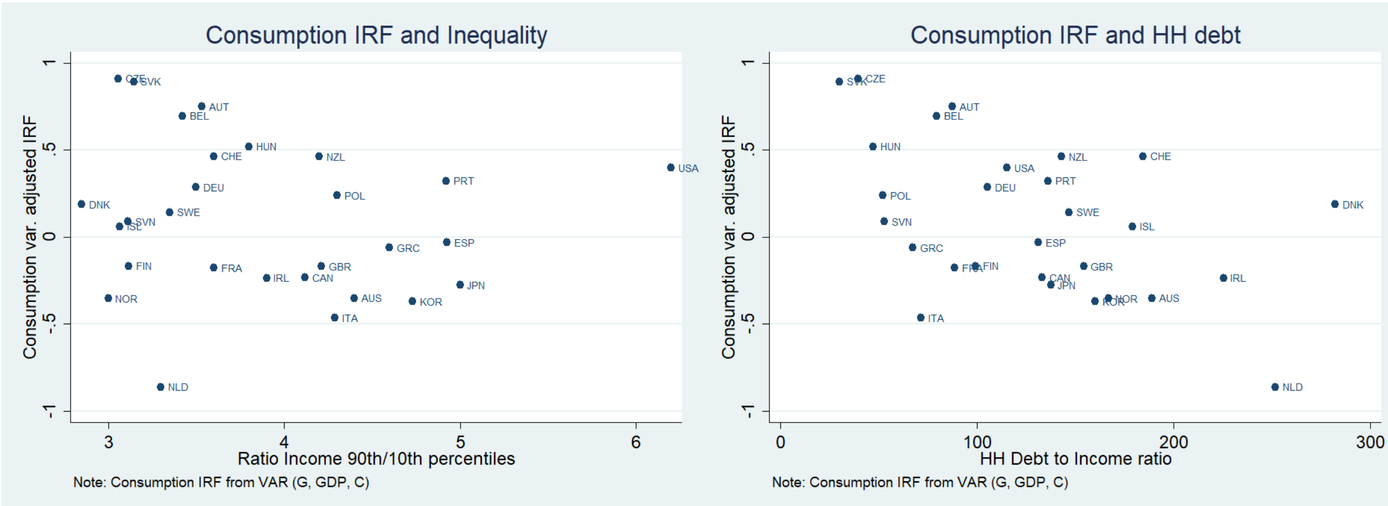


Figure 3
Consumption Response and Debt-burdened

## 4.2 U.S. Cross-County consumption response and debt

The cross-country evidence above is somewhat surprising given recent empirical work that finds that high household leverage is associated with higher rather than lower consumption responses to government spending. In particular, Demyanyk et al. (2016) demonstrate that during the great recession, an increase in government spending in a region was associated with a consumption response among households in that region that was increasing in leverage. High debt, they concluded, was associated with high MPCs. In contrast, our cross-county regressions imply that high debt is associated with low MPCs.

To reconcile these findings, it is important to note that the Demyanyk et al. (2016) study was based on the Great Recession, when the supply of credit was limited (see for example, Mian and Sufi (2015)). In our more general framework with minimum consumption thresholds (see Miranda-Pinto et al. (2018)), a contraction in credit supply can cause high-debt households to be unable to afford even their minimum level of consumption and render them credit-constrained (rather than saving-constrained). During normal times (with greater credit supply), high debt is instead associated with saving constraints and low MPCs. Indeed, Demyanyk et al. (2016) presented evidence that government spending multipliers were increasing in debt during the Great Recession but not during the boom period of the mid-2000s. One possible explanation for their findings is that consumption multipliers were lower (or at least not higher) in high-debt areas prior to the Great Recession.

Here we test this hypothesis directly using data across counties in the United States. Our measure of consumption is auto registrations, which has been used as a proxy for broad measures of consumption in cross-sectional analyses of disaggregate levels of economic geography such as counties (e.g., Demyanyk et al. (2016), Mian et al. (2013)). The data are

Table 3
Consumption Response and Country Characteristics

| VARIABLES | (1) CIRF | (2) CIRF | (3) CIRF |
|---|---|---|---|
| HH debt to income | -0.219** | -0.220** | -0.248*** |
| | (0.025) | (0.026) | (0.005) |
| Inflation Targeting | | -0.088 | |
| | | (0.548) | |
| Fiscal Mult. 4 qtrs | | | 0.092 |
| | | | (0.379) |
| Observations | 28 | 28 | 28 |
| R-squared | 0.249 | 0.259 | 0.289 |

* Note: This table presents the OLS coefficients of regressing the estimated CIRF – variance adjusted – response of private consumption to government innovations against household debt (from OECD database), inflation targeting dummy (Carare and Stone (2003)), and the fiscal multiplier. The fiscal multiplier is estimated analogously to the IRRF from the Blanchard and Perotti (2002)'s approach. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

provided by R. L. Polk. The government spending measure is based on the DOD spending measure from Demyanyk et al. (2016), which begins in 2001 (see also Auerbach et al. (2018)). Our measure of county-level debt-to-income, which spans from 2001 through 2007, is from Mian and Sufi (2015).

Our empirical specification is

$$\frac{C_{i,t} - C_{i,t-1}}{C_{i,t-1}} = \beta_0 \frac{G_{i,t} - G_{i,t-1}}{Y_{i,t-1}} + \beta_1 \frac{G_{i,t} - G_{i,t-1}}{Y_{i,t-1}} \cdot DTI_{i,t-1} + \gamma DTI_{i,t-1} + \alpha_i + \lambda_t + \epsilon_{i,t}$$

where $Y_{i,t}$ is income in county $i$ in year $t$, $C$ is auto registrations, $DTI$ is household leverage, $G$ is military spending, and $\alpha_i$ and $\lambda_t$ are location and time fixed effects. The coefficient of interest is $\beta_1$, which is an estimate of the extent to which the consumption response to fiscal stimulus depends on households leverage. We instrument for the change in defense spending (and its interaction with leverage) using the Bartik-type instrument used in Nakamura and Steinsson (2014), Demyanyk et al. (2016), and Auerbach et al. (2018). Specifically, $(\frac{G_{i,t}-G_{i,t-1}}{Y_{i,t-1}})$ is instrumented with $(s_i \cdot \frac{G_t - G_{t-1}}{Y_{l,t-1})})$ where $G_t$ is aggregate government spending and $s_i$ is the average share of county $i$ in total government spending over the sample period. This IV approach addresses two potential concerns. First, as discussed in Nakamura and Steinsson (2014), it corrects for the possibility that defense spending may respond endogenously to local economic conditions. Second, the instrument captures the component of defense contracts that represent actual spending/production increases (see the discussion in Auerbach et al. (2018)).

Our specification is most similar to Demyanyk et al. (2016) in that it includes the interaction between defense spending and debt. It differs in a couple of important respects. First, ours is a panel specification, which allows us to absorb county-specific factors in fixed effects. Second, we focus on pre-recession (2001 through 2007) data using county-level data rather than city-level data. Conducting the analysis at the county level provides more cross-sectional variation and a more precise estimate of the debt-dependence of consumption responses to defense spending during periods of normal-to-high credit supply. Third, our dependent variable is the percentage change in consumption (rather than the change normalized by lagged income), which implies that the coefficients on government spending should be interpreted as an elasticity (rather than a multiplier). Since we do not know the value of the automobiles registered, using a percent change is more natural than trying to infer auto values to derive a specific consumption multiplier. That being said, the results we present below are qualitatively similar when normalizing the change in consumption by lagged local income rather than lagged local consumption

Table 4 shows that the response of auto purchases to local defense spending is indeed lower in counties with higher debt (columns 2 and 3). While the direct response of auto purchases appears negligible (column 1), this measure of the average effects masks heterogeneity due to household leverage. Counties that have higher leverage have a smaller response of auto purchases to government spending.[15] Our evidence from auto purchases is consistent with the evidence in Demyanyk et al. (2016) that fiscal multipliers were, if anything, smaller in high-debt regions during the mid-2000s.

## 4.3 Credit constrained vs. saving constrained households

In this final section, we discuss in more detail the implications of credit constrained and saving constrained households for the response of the macroeconomy to fiscal shocks. We start by discussing what are the implications of credit constrained households for the response of interest rates (and consumption) to fiscal shocks. We then directly compare our empirical results to Brinca et al. (2016), who emphasize the role of credit constraints in explaining the cross-country relationship between fiscal multipliers and inequality.

As we demonstrate in our Appendix B, credit constraints are not able to account for our cross-country facts for the response of interest rates to fiscal shocks. We modify our model in section 3 and assume that rather than saving constrained, poor households are credit constrained. If credit conditions are sufficiently tight, then poor households cannot

[15] We have run similar specifications using city-level data. While the statistical significance of the interaction term varies across specifications, each similarly exhibits consumption responses that are decreasing rather than increasing in debt in the pre-recession period.

Table 4
Consumption and Household Debt US counties

| VARIABLES | (1) $\%\Delta C$ | (2) $\%\Delta C$ | (3) $\%\Delta C$ |
|---|---|---|---|
| $\%\Delta G$ | 0.19 | | 4.28*** |
| | (0.80) | | (1.23) |
| $DTI$ | | -0.05 | -0.03 |
| | | (0.08) | (0.08) |
| $\%\Delta G \cdot DTI$ | | | -3.58*** |
| | | | (1.22) |
| Observations | 8286 | 8286 | 8286 |
| First-stage F-stat | 7.14 | | 8.40 |

Note: This table presents the coefficients of regressing the percent change in county $i$'s auto registrations ($\%\Delta C$) against household debt to income in county $i$ (DTI) from Mian and Sufi (2015), the percent change in defense spending in county $i$ ($\%\Delta G$), and the interaction between these two covariates. We instrument for the change in defense spending (and its interaction with leverage) using the Bartik-type instrument used in Nakamura and Steinsson (2014), Demyanyk et al. (2016), and Auerbach et al. (2018). Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

afford their minimum consumption threshold and they exhibit high MPCs. In this case the interest rate response to fiscal shocks is increasing in the fraction of poor credit-constrained households. Thus, a model with tight credit conditions predicts a positive relationship between the IRRF and inequality (household debt), contrary to what we document for the group of OECD economies before the GFC.

Accounting for variation in credit conditions can help reconcile our findings with those in Brinca et al. (2016). The authors show that for a group of 31 OECD and non-OECD countries, there is a positive correlation between inequality (wealth inequality) and fiscal multipliers. They explain the facts with a model of uninsurable idiosyncratic risk where borrowing constraints and precautionary savings play a key role.

Here we show that for the group of OECD countries and for the period before the GFC, fiscal multipliers are not related to inequality or household debt as predicted by a model with credit constrained households. Figure 4 shows a scatter plot of our GDP four quarter cumulative response to government spending shocks against income inequality, using Blanchard and Perotti (2002) approach. If anything, we observe a negative relationship between our measure of fiscal multiplier and income inequality.

Similarly, Figure 5 depicts the relationship between our fiscal multiplier and household debt. We do not observe the positive relationship that would be consistent with hand-to-mouth agents.

Hence, we conclude that while the results in Brinca et al. (2016) can be useful to un-

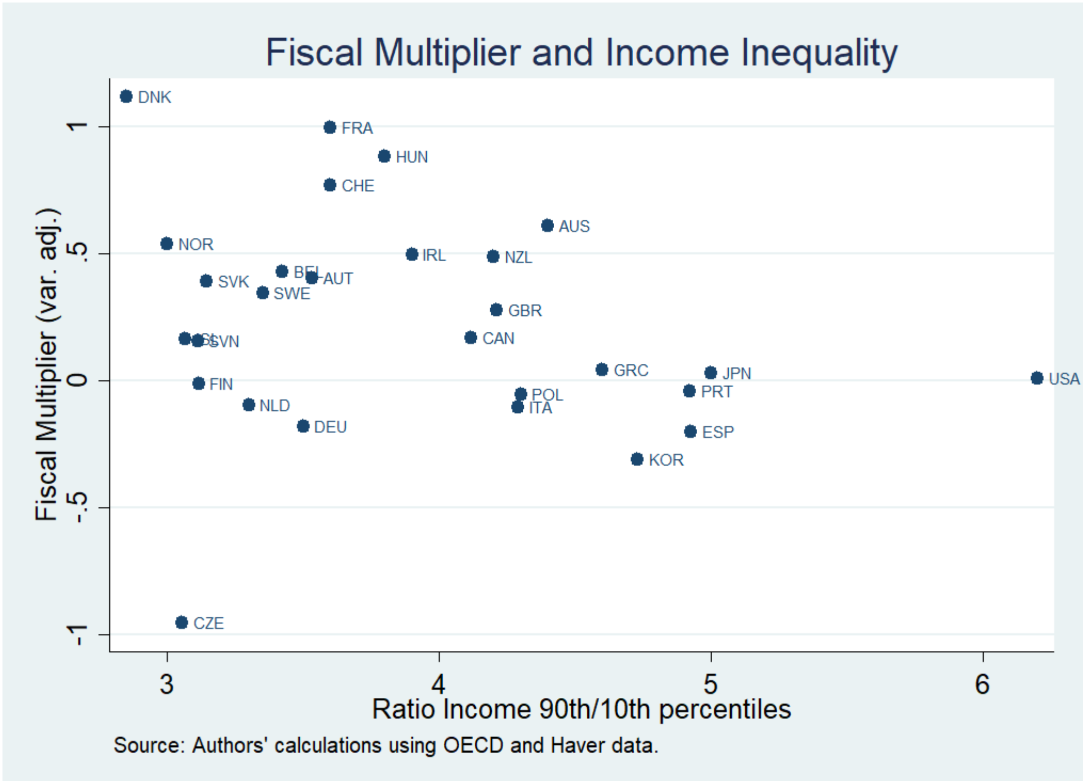


Figure 4
GDP response to fiscal shocks and Inequality

derstand the heterogeneity in fiscal multipliers between OECD and non-OECD countries, especially when credit conditions are tight, an alternative framework is necessary for understanding fiscal transmission mechanisms in OECD countries when credit conditions are relatively loose.

# 5 Conclusions

We present new evidence that, during the years before the Global Financial Crisis, the effect of government spending on interest rates (IRRF) varies across countries, with half of OECD countries exhibiting a negative interest rate response. The IRRF is decreasing in country-level inequality (or household debt), contrary to the predictions of existing heterogeneous agent models with credit constraints.

We interpret this evidence through the lens of a theoretical framework in which the interest rate response to demand stimulus depends on the share of consumers who are low-wealth and burdened with debt due to saving constraints (minimum consumption thresholds). In our setting, debt burdens do not reflect credit constraints but rather result from households' minimum consumption needs. This additional debt is burdensome in the sense that households pay it off more quickly out of additional income than they would in the absence of a

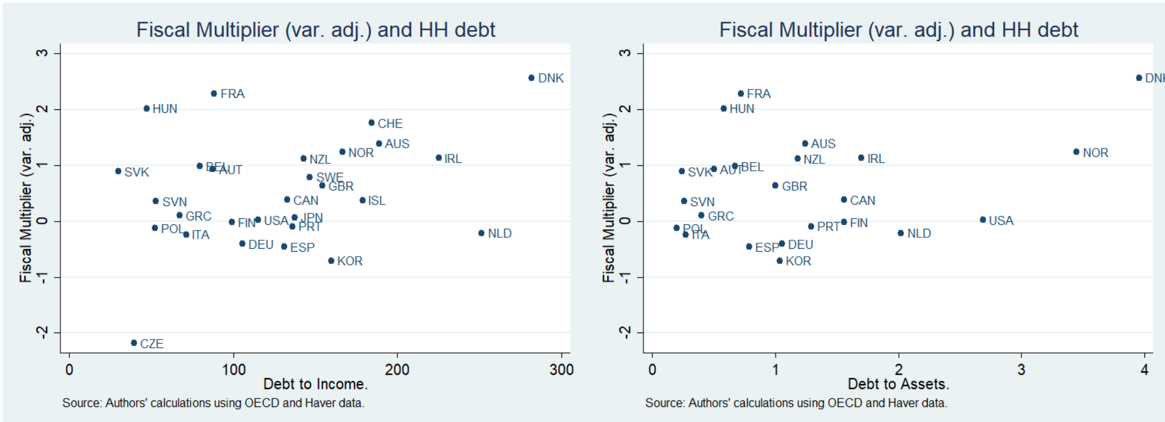


Figure 5
GDP response to fiscal shocks and Household Debt

minimum consumption constraint.

In Miranda-Pinto et al. (2018), we document that saving constraints can also rationalize otherwise unexplained features of the microdata. Here we demonstrate the relevance of saving constraints for the effects of macroeconomic shocks. In particular, our model with saving-constrained households can rationalize why the interest rate and consumption responses to government purchases are declining in inequality and household debt. A key implicaion of our findings is that not only does the effect of fiscal stimulus depend on debt, as has been documented in recent empirical and theoretical work, but that the sign of debt-dependence varies with credit conditions. When credit is loose and poor households can borrow to meet their minimum consumption thresholds, fiscal stimulus can redistritute resources to low-MPC poor households and relax credit markets. When credit conditions are tight, these poor households are credit-constrained and have high MPCs.

# References

Aiyagari, S. R. (1994). Uninsured idiosyncratic risk and aggregate saving. *Quarterly Journal of Economics*, 109(3):569–684.

Auerbach, A. and Gorodnichenko, Y. (2012). Measuring the output responses to fiscal policy. *American Economic Journal: Economic Policy*, 4(2):1–27.

Auerbach, A. and Gorodnichenko, Y. (2013). Output spillovers from fiscal policy. *American Economic Review: Papers and Proceedings*, 103(3).

Auerbach, A., Gorodnichenko, Y., and Murphy, D. (2018). Local fiscal multipliers and fiscal spillovers in the united states. *mimeo*.

Bachmann, R. and Sims, E. R. (2012). Confidence and the transmission of government spending shocks. *Journal of Monetary Economics*, 59(3):235–249.

Barro, R. J. (1987). Government spending, interest rates, prices, and budget deficits in the United Kingdom, 1701-1918. *Journal of Monetary Economics*, 20(2):221–247.

Blanchard, O. and Perotti, R. (2002). An empirical characterization of the dynamic effects of changes in government spending and taxes on output*. *The Quarterly Journal of Economics*, 117(4):1329.

Brinca, P., Holter, H. A., Krusell, P., and Malafry, L. (2016). Fiscal mutipliers in the 21st century. *Journal of Monetary Economics*, 77:53–69.

Broner, F., Clancy, D., Erce, A., and Martin, A. (2018). Fiscal multipliers and foreign holdings of public debt. *Working Paper*.

Carare, A. and Stone, M. R. (2003). Inflation targeting regimes. Working papers, IMF.

Chetty, R. and Szeidl, A. (2007). Consumption commitments and risk preferences. *Quarterly Journal of Economics*, 122(2):831–877.

Corsetti, G., Meier, A., and Mueller, G. J. (2012a). Fiscal Stimulus with Spending Reversals. *The Review of Economics and Statistics*, 94(4):878–895.

Corsetti, G., Meier, A., and Mueller, G. J. (2012b). What determines government spending multipliers? *Economic Policy*, 27(72):521–565.

Demyanyk, Y., Loutskina, E., and Murphy, D. (2016). Fiscal Stimulus and Consumer Debt.

Eggertsson, G. B. and Krugman, P. (2012). Debt, Deleveraging, and the Liquidity Trap: A Fisher-Minsky-Koo Approach. *The Quarterly Journal of Economics*, 127(3):1469–1513.

Faccini, R., Mumtaz, H., and Surico, P. (2016). International fiscal spillovers. *Journal of International Economics*, 99(C):31–45.

Fisher, J. and Peters, R. (2010). Using Stock Returns to Identify Government Spending Shocks. *Economic Journal*, 120(544):414–436.

Huggett, M. (1993). The risk-free rate in heterogeneous-agent incomplete-insurance economies. *Journal of Economic Dynamics and Control*, 17:953 – 969.

Ilzetzki, E., Mendoza, E. G., and Vegh, C. A. (2013). How big (small?) are fiscal multipliers? *Journal of Monetary Economics*, 60(2):239–254.

Jappelli, T. and Pistaferri, L. (2014). Fiscal policy and mpc heterogeneity. *American Economic Journal: Macroeconomics*, 6(4):107–136.

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

Mian, A., Rao, K., and Sufi, A. (2013). Household balance sheets, consumption, and the economic slump. *Quarterly Journal of Economics*, 128(4):1687–1726.

Mian, A. and Sufi, A. (2015). What explains the 2007-2009 drop in employment? *Econometrica*, 82(6),:2197–2223.

Miranda-Pinto, J., Murphy, D., Walsh, K., and Young, E. (2018). Saving-Constrained Households. *Mimeo University of Virginia.*

Misra, K. and Surico, P. (2014). Consumption, Income Changes, and Heterogeneity: Evidence from Two Fiscal Stimulus Programs. *American Economic Journal: Macroeconomics*, 6(4):84–106.

Murphy, D. and Walsh, K. (2017). Government Spending and Interest Rates. *Mimeo University of Virginia.*

Nakamura, E. and Steinsson, J. (2014). Fiscal stimulus in a monetary union: Evidence from u.s. regions. *American Economic Review*, 104(3):753–792.

Priftis, R. and Zimic, S. (2018). Sources of borrowing and fiscal multipliers. *Bank of Canada Staff Working Paper*, (2018-32).

Ramey, V. A. (2011). Identifying Government Spending Shocks: It's all in the Timing. *The Quarterly Journal of Economics*, 126(1):1–50.

Ravn, M. O., Schmitt-Groh?, S., and Uribe, M. (2012). Consumption, government spending, and the real exchange rate. *Journal of Monetary Economics*, 59(3):215–234.

Rossi, B. and Zubairy, S. (2011). What Is the Importance of Monetary and Fiscal Shocks in Explaining U.S. Macroeconomic Fluctuations? *Journal of Money, Credit and Banking*, 43(6):1247–1270.

Sahm, C. R., Shapiro, M. D., and Slemrod, J. (2015). Balance-Sheet Households and Fiscal Stimulus: Lessons from the Payroll Tax Cut and Its Expiration. NBER Working Papers 21220, National Bureau of Economic Research, Inc.

# Appendix A: Robustness Checks, Tables, and Figures

| | G | | | GDP | | Interest Rates Haver | | | | C | |
|---|---|---|---|---|---|---|---|---|---|---|---|
| **Country** | OECD | Haver | Haver | OECD | Haver | OECD sr | Tbill Haver | Policy Rate | G bond Haver | OECD | Haver |
| Australia | 1959-Q4 | 1959-Q3 | 1959-Q3 | 1959-Q3 | 1959-Q3 | 1968-Q1 | 1969-Q3 | 1969-Q3 | 1957-Q1 | 1959-Q3 | - |
| Austria | 1988-Q2 | 1957-Q1 | 1999-Q1 | 1988-Q1 | 1999-Q1 | 1989-Q3 | - | 1957-Q1 | 1971-Q1 | 1996-Q1 | 1996-Q1 |
| Belgium | 1995-Q1 | 1980-Q1 | 1999-Q1 | 1995-Q1 | 1999-Q1 | 1958-Q1 | 1957-Q1 | 1957-Q1 | 1957-Q1 | 1995-Q1 | 1995-Q1 |
| Canada | 1961-Q2 | 1957-Q1 | 1957-Q1 | 1961-Q1 | 1957-Q1 | 1956-Q1 | 1957-Q1 | 1992-Q4 | 1957-Q1 | 1982-Q1 | 1957-Q1 |
| Czech Republic | 1996-Q2 | 1990-Q1 | 1990-Q1 | 1996-Q1 | 1990-Q1 | 1993-Q1 | 1993-Q4 | 1995-Q4 | 2000-Q2 | 1996-Q1 | 1995-Q1 |
| Denmark | 1995-Q1 | 1977-Q1 | 1977-Q1 | 1995-Q1 | 1977-Q1 | 1987-Q1 | - | 1957-Q1 | 1960-Q1 | 1995-Q1 | 1995-Q1 |
| Finland | 1990-Q1 | 1970-Q1 | 1988-Q1 | 1990-Q1 | 1999-Q1 | 1987-Q1 | - | 1957-Q1 | 1988-Q1 | 1990-Q1 | 1990-Q1 |
| France | 1955-Q1 | 1999-Q1 | 1999-Q1 | 1955-Q1 | 1999-Q1 | 1955-Q2 | 1970-Q1 | 1964Q-1 | 1957-Q1 | 1980-Q1 | 1980-Q1 |
| Germany | 1970-Q1 | 1999-Q1 | 1999-Q1 | 1970-Q1 | 1999-Q1 | 1960-Q1 | 1975-Q1 | 1957-Q1 | 1957-Q1 | 1991-Q1 | 1991-Q1 |
| Greece | 1970-Q2 | - | - | 1970-Q1 | 2000-Q4 | 1980-Q2 | 1985-Q2 | 1957-Q1 | 1992-2016 | 1995-Q1 | 1995-Q1 |
| Hungary | 1995-Q1 | 1995-Q1 | 1995-Q1 | 1995-Q1 | 1995-Q1 | 1992-Q2 | 1989-Q2 | 1985-Q1 | 2001-Q1 | 1995-Q1 | 1995-Q1 |
| Iceland | 1996-Q4 | - | 1997-Q1 | 1997-Q1 | 1997-Q1 | 1988-Q1 | 1987-Q3 | 1964-Q1 | 1992-Q1 | 1997-Q1 | - |
| Ireland | 1996-Q5 | 1997-Q1 | 1999-Q1 | 1997-Q1 | 1999-Q1 | 1984-Q1 | 1973Q1 | 195Q17 | 1964-Q1 | 1995-Q1 | 1995-Q1 |
| Italy | 1981-Q1 | 1999-Q1 | 1999-Q1 | 1981-Q4 | 1999-Q1 | 1978-Q4 | 1977-Q2 | 1964-Q1 | 1958-Q1 | 1996-Q1 | 1995-Q1 |
| Japan | 1994-Q1 | 1957-Q1 | 1957-Q1 | 1994-Q1 | 1957-Q1 | 1985-Q3 | 1957-Q1 | 1957-Q1 | 1966-Q1 | 1994-Q1 | 1957-Q1 |
| Korea | 1970-Q1 | 1960-Q1 | 1960-Q1 | 1970-Q1 | 1960-Q1 | 1991-Q1 | - | 1999-Q2 | 1973-Q2 | 1960-Q1 | 2001-Q1 |
| Netherlands | 1988-Q1 | 1990-Q1 | 1990-Q1 | 1988-Q1 | 1990-Q1 | 1986-Q1 | 1978-Q2 | - | 1965-Q1 | 1996-Q1 | 1995-Q1 |
| New Zealand | 1987-Q1 | 1987-Q2 | 1987-Q2 | 1987-Q2 | 1987-Q2 | 1974-Q1 | 1978-Q1 | 1999-Q1 | 1964-Q1 | 1987-Q2 | 1987-Q2 |
| Norway | 1978-Q1 | 1961-Q1 | 1961-Q1 | 1978-Q1 | 1961-Q1 | 1979-Q1 | - | 1964-Q1 | 1961-Q4 | 1978-Q1 | 1978-Q1 |
| Poland | 1995-Q1 | 1995-Q1 | 1995-Q1 | 1995-Q1 | 1995-Q1 | 1991-Q3 | 1992-Q1 | 1998-Q1 | 2001-Q1 | 1995-Q1 | 1995-Q1 |
| Portugal | 1995-Q1 | 1999-Q1 | 1999-Q1 | 1995-Q1 | 1999-Q1 | 1985-Q4 | 1985-Q4 | 1957Q1 | 1957-Q1 | 1995-Q1 | 1995-Q1 |
| Slovak Republic | 1997-Q1 | 1993-Q1 | - | 1997-Q1 | 2009-Q1 | 1995-Q3 | - | 2001-Q2 | 2000-q4 | 1995-Q1 | - |
| Slovenia | 1995-Q1 | 1995-Q1 | 2007-Q1 | 1995-Q1 | 2007-Q1 | 2002-Q1 | 1998-Q3 | 1992-Q1 | 1991-Q4 | 1995-Q1 | 1995-Q1 |
| Spain | 1995-Q1 | 1999-Q1 | 1999-Q1 | 1995-Q1 | 1999-Q1 | 1977-Q1 | 1979-Q1 | 1964-Q1 | 1978-Q2 | 1995-Q1 | 1995-Q1 |
| Sweden | 1960-Q1 | - | 1980-Q1 | 1960-Q1 | 1980-Q1 | 1955-Q1 | 1963-Q2 | 2002-Q3 | 1960-Q1 | 1993-Q1 | 1993-Q1 |
| Switzerland | 1980-Q1 | 1970-Q1 | 1970-Q1 | 1980-Q1 | 1970-Q1 | 1974-Q1 | 1980-Q1 | 1964-Q1 | 1964-Q1 | 1980-Q1 | 1970-Q1 |
| United Kingdom | 1955-Q1 | 1957-Q1 | 1957-Q1 | 1955-Q1 | 1957-Q1 | 1978-Q1 | 1964-Q1 | 1959-Q1 | 1957-Q1 | 1995-Q1 | 1957-Q1 |
| United States | 1955-Q1 | 1957-Q1 | 1957-Q1 | 1955-Q1 | 1957-Q1 | 1955-Q1 | 1957-Q1 | 1982-Q3 | 1957-Q1 | 1955-Q1 | 1957-Q1 |

Figure 6
Sample for VAR estimation (Green means haver is used)

## Auerbach and Gorodnichenko (2013) shocks and local projection methods

Here we use the government spending shocks estimated by Auerbach and Gorodnichenko (2013). The authors regress one-period-ahead percent forecast errors for government spending from the OECD's "Outlook and Projections Database" in each country on that country's lagged macroeconomic variables (output, government spending, exchange rate, inflation, investment, and imports). The authors also consider a set of country and period fixed-effects. The residual from this regression are innovations in government spending orthogonal to professional forecasts and lags of macroeconomic variables.[16]

We take the estimated unanticipated government spending shocks from Auerbach and Gorodnichenko (2013) and use linear projection methods to measure the effect on government bond yields. The data is semi-annual, therefore, to match our four quarters cumulative IRRF from section 2.1, we regress the semi-annual government bond yield against the contemporaneous innovation to government spending and its one semester lag. In particular, for each country, we regress

$$r_t = \beta_0 + \beta_1 \hat{G}_t^{shock} + \beta_2 \hat{G}_{t-1}^{shock} + \mu_t, \tag{5}$$

where $r_t$ is the country's government bond yield at semester $t$ and $\hat{G}_t^{shock}$ is the Auerbach and Gorodnichenko (2013) semi-annual identified shock to government spending in semester $t$. We convert our quarterly data on government bond yields to semi-annual frequency by simply taking the average of each semester's quarters. The cumulative 4 quarters (2 semesters) is simply $IRRF = \hat{\beta}_1 + \hat{\beta}_2$. We use the OLS (one) standard deviation of the coefficients ($sd(\hat{\beta})$) to adjust for uncertainty in the estimates. In particular, we define

$$IRRF = \frac{\hat{\beta}_1 + \hat{\beta}_2}{[\hat{\beta}_1 + sd(\hat{\beta}_1) + \hat{\beta}_2 + sd(\hat{\beta}_2)] - [\hat{\beta}_1 - sd(\hat{\beta}_1) + \hat{\beta}_2 - sd(\hat{\beta}_2)]}$$

Figure 7 reports the estimated variance adjusted IRRF using this approach. There are 13 countries with negative IRRF. Surprisingly, the U.S. displays a positive and precise positive IRRF. The key difference with respect to the IRRF for the U.S obtained using the approach in Blanchard and Perotti (2002) is that in this case we have a significantly smaller amount of observations. Indeed, we only have government spending shocks identified semi-annually since 1986 semester 1, while in the Blanchard and Perotti (2002) approach we have quarterly data since 1957Q1. Greece is another country with significant differences across methods. Greece displays the most negative IRRF using the local projection method, while it has an almost zero IRRF using the Blanchard and Perotti (2002) approach. These results are also a

[16] Note that the government spending series in Auerbach and Gorodnichenko (2013) is the sum of real public consumption expenditure and real government gross capital formation.

consequence of the small sample size. With local projection method we have Greece's shocks data since 1997 semester 1 until 2003 semester 2, while for the Blanchard and Perotti (2002) approach we have quarterly data for the period 1992-2007.

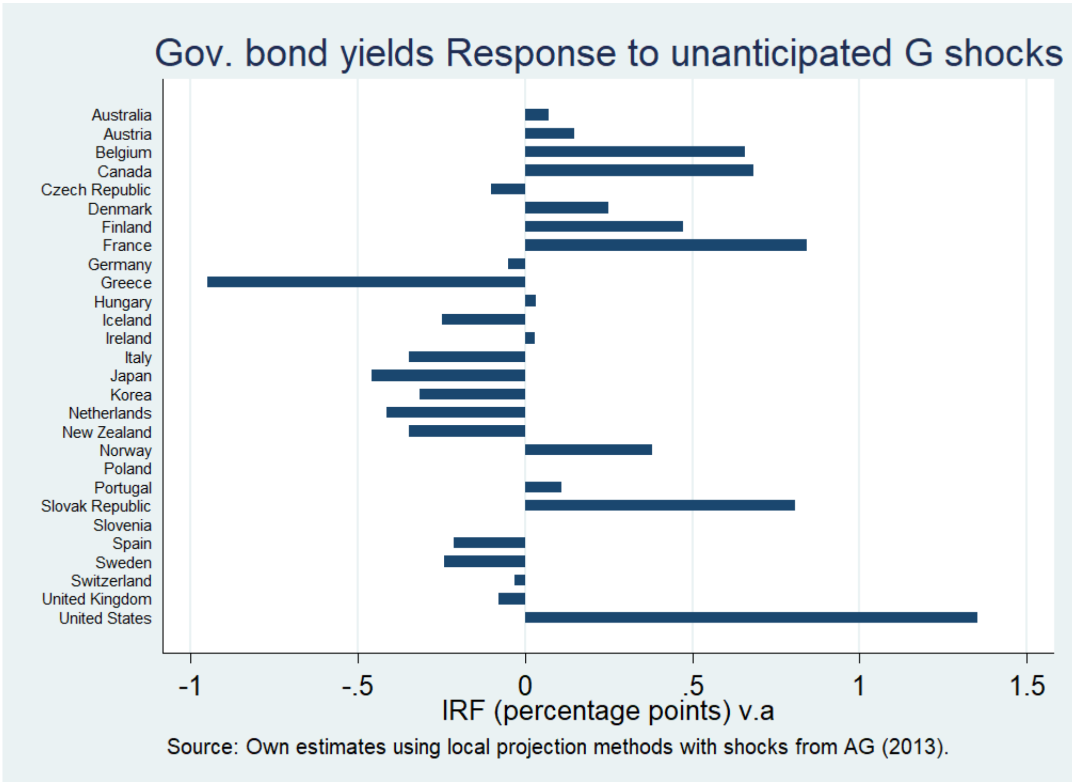


Figure 7
Effect of fiscal shocks on government bond yields across OECD Countries Auerbach and Gorodnichenko (2013) shocks

We now turn our attention to the identified shocks using local projection methods and the Auerbach and Gorodnichenko (2013) shocks. In Figure 8, we show that the same relationship between IRRF and inequality holds when we use local projection methods and semi-annual government innovations from Auerbach and Gorodnichenko (2013). Auerbach and Gorodnichenko (2013) account for the potential anticipation of government spending shocks by using OECD data on professional forecasters, and lags of macroeconomics variables. We can observe that the U.S and Greece are indeed two outliers in our data. Hence, we proceed to drop these two countries from our regression analysis.

In Tables 5-6 we report the OLS estimation of regressing IRRF against country characteristics. The results are very similar to the obtained in Tables 1 and 2. A one standard deviation increase in income inequality is associated with a 18 basis point decline in the IRRF. This result is robust to adding additional country characteristics. The results for households debt are less precise but also suggest that a one standard deviation increase in income inequality is associated with approximately a 18 basis point decline in the IRRF (column 4 of Table 6 ). Similar results hold when we use the point estimate – not variance adjusted – of the IRRF in Tables 10-11 of our Appendix.

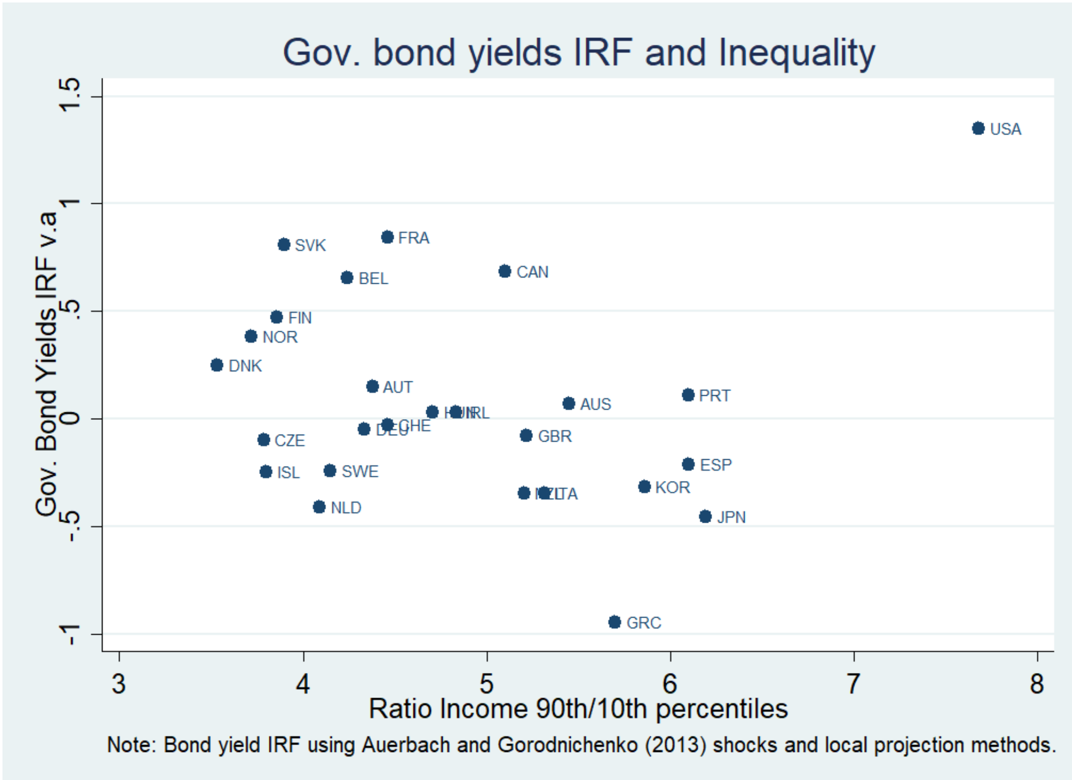


Figure 8
Inequality and Gov. Bond Yield Response Auerbach and Gorodnichenko (2013)'s shocks

Table 5
Gov. Bond Yields Response and Country Characteristics (Auerbach and Gorodnichenko (2013)'s shocks)

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| VARIABLES | IRRF | IRRF | IRRF | IRRF | IRRF |
| Income ratio 90th/10th | -0.180** | -0.182** | -0.176** | -0.144* | -0.091 |
| | (0.077) | (0.079) | (0.077) | (0.077) | (0.107) |
| Financial Openness | | -0.013 | | | |
| | | (0.045) | | | |
| Inflation Targeting | | | -0.105 | | |
| | | | (0.152) | | |
| Fiscal Mult. 4 qtrs | | | | 0.130** | |
| | | | | (0.056) | |
| Ext. G Debt to GDP | | | | | 0.044 |
| | | | | | (0.056) |
| Observations | 24 | 24 | 24 | 24 | 17 |
| R-squared | 0.144 | 0.145 | 0.163 | 0.255 | 0.051 |

* Note: This table presents the OLS coefficients of regressing the estimated variance adjusted two semester cumulative response of government bond yields to government innovations against income inequality (from OECD database), financial openness (from Lane and Milesi-Ferretti (2007)), inflation targeting dummy (Carare and Stone (2003)), foreign government debt to GDP (from IMF-World Bank QPSD data), and the fiscal multiplier. The fiscal multiplier is estimated analogously to the IRRF from the Blanchard and Perotti (2002)'s approach. We drop the Greece and the U.S. from the sample. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Table 6
Gov. Bond Yields Response and Country Characteristics (Auerbach and Gorodnichenko (2013)'s shocks)

| VARIABLES | (1) IRRF | (2) IRRF | (3) IRRF | (4) IRRF |
|---|---|---|---|---|
| HH debt to income | -0.119 | -0.146 | -0.123 | -0.175** |
| | (0.082) | (0.100) | (0.079) | (0.070) |
| Financial Openness | | 0.068 | | |
| | | (0.060) | | |
| Inflation Targeting | | | -0.136 | |
| | | | (0.159) | |
| Fiscal Mult. 4 qtrs | | | | 0.202*** |
| | | | | (0.064) |
| Observations | 24 | 24 | 24 | 24 |
| R-squared | 0.095 | 0.122 | 0.127 | 0.359 |

* Note: This table presents the OLS coefficients of regressing the estimated variance adjusted two semester cumulative response of government bond yields to government innovations against income inequality (from OECD database), financial openness (from Lane and Milesi-Ferretti (2007)), inflation targeting dummy (Carare and Stone (2003)), and the fiscal multiplier. The fiscal multiplier is estimated analogously to the IRRF from the Blanchard and Perotti (2002)'s approach. We drop Greece and the U.S from the sample. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

## Policy rate response to fiscal shocks and inequality

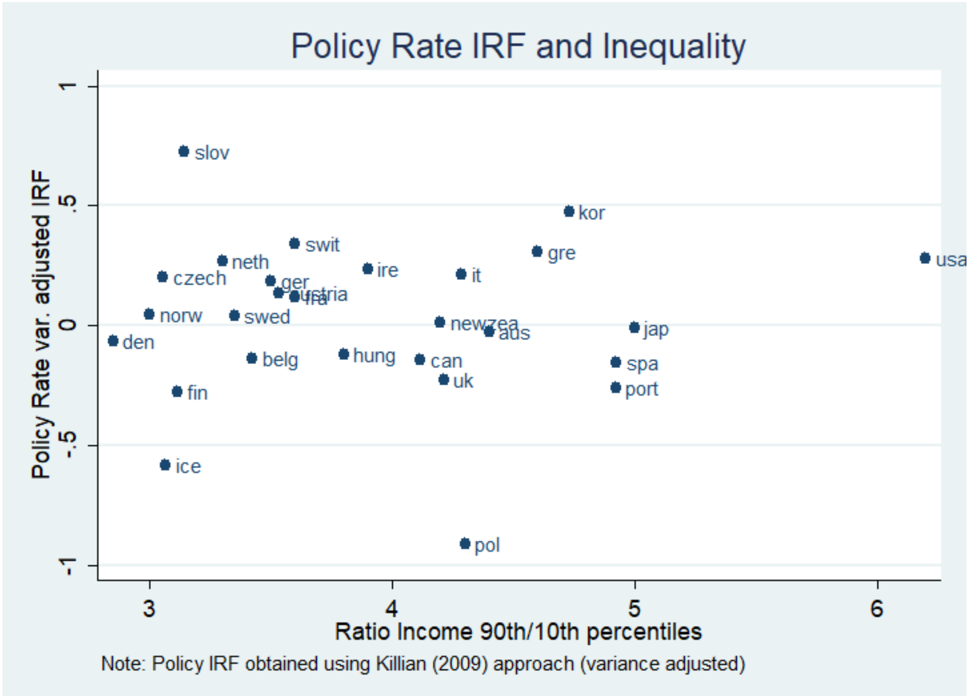


Figure 9
Inequality and Policy Rates Response

## Not-variance adjusted IRRF and CIRF

Table 7
IRRF not var. adj. and Inequality (Blanchard and Perotti (2002))

| VARIABLES | (1) IRRF | (2) IRRF | (3) IRRF | (4) IRRF | (5) IRRF |
|---|---|---|---|---|---|
| Income ratio 90th/10th | -0.112* | -0.103 | -0.116** | -0.096* | -0.154* |
| | (0.059) | (0.063) | (0.056) | (0.054) | (0.078) |
| Financial Openness | | 0.049 | | | |
| | | (0.050) | | | |
| Inflation Targeting | | | -0.212* | | |
| | | | (0.122) | | |
| Fiscal Mult. 4 qtrs | | | | 0.064 | |
| | | | | (0.058) | |
| Ext. G Debt to GDP | | | | | -0.025 |
| | | | | | (0.061) |
| Observations | 28 | 28 | 28 | 28 | 19 |
| R-squared | 0.106 | 0.125 | 0.204 | 0.139 | 0.197 |

* Note: This table presents the OLS coefficients of regressing the estimated four quarters cumulative response of government bond yields to government spending shocks against income inequality (from OECD database), financial openness (from Lane and Milesi-Ferretti (2007)), inflation targeting dummy (Carare and Stone (2003)), foreign government debt to GDP (from IMF-World Bank QPSD data), and the fiscal multiplier. The fiscal multiplier is estimated analogously to the IRRF. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Table 8

IRRF not var. adj. and Household Debt (Blanchard and Perotti (2002))

| VARIABLES | (1)<br>IRRF | (2)<br>IRRF | (3)<br>IRRF | (4)<br>IRRF |
|---|---|---|---|---|
| HH debt to assets | -0.110* | -0.118* | -0.109* | -0.136* |
| | (0.058) | (0.059) | (0.059) | (0.068) |
| Financial Openness | | 0.052 | | |
| | | (0.036) | | |
| Inflation Targeting | | | -0.149 | |
| | | | (0.142) | |
| Fiscal Mult. 4 qtrs | | | | 0.097 |
| | | | | (0.084) |
| Observations | 23 | 23 | 23 | 23 |
| R-squared | 0.096 | 0.118 | 0.141 | 0.151 |

* Note: This table presents the OLS coefficients of regressing the estimated four quarters cumulative response of government bond yields to government spending shocks against median household debt to assets ratio (from OECD database), financial openness (from Lane and Milesi-Ferretti (2007)), inflation targeting dummy (Carare and Stone (2003)), and the fiscal multiplier. The fiscal multiplier is estimated analogously to the IRRF. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Table 9
Consumption Response (not-variance adjusted) Blanchard and Perotti (2002) and Country Characteristics

| VARIABLES | (1)<br>CIRF | (2)<br>CIRF |
|---|---|---|
| HH debt to income | -0.005** | -0.005*** |
| | (0.029) | (0.005) |
| Inflation Targeting | -0.000 | |
| | (0.911) | |
| Fiscal Mult. 4 qtrs | | 0.002 |
| | | (0.268) |
| Observations | 28 | 28 |
| R-squared | 0.274 | 0.337 |

[*] Note: This table presents the OLS coefficients of regressing the estimated CIRF response of private consumption to government innovations against household debt (from OECD database), inflation targeting dummy (Carare and Stone (2003)), and the fiscal multiplier. The fiscal multiplier is estimated analogously to the IRRF from the Blanchard and Perotti (2002)'s approach. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Table 10

IRRF not var. adj. and Inequality Auerbach and Gorodnichenko (2013)'s shocks

| VARIABLES | (1) IRRF | (2) IRRF | (3) IRRF | (4) IRRF | (5) IRRF |
|---|---|---|---|---|---|
| Income ratio 90th/10th | -0.284** | -0.282** | -0.280** | -0.201* | -0.161 |
| | (0.114) | (0.112) | (0.115) | (0.105) | (0.141) |
| Financial Openness | | 0.010 | | | |
| | | (0.059) | | | |
| Inflation Targeting | | | -0.106 | | |
| | | | (0.298) | | |
| Fiscal Mult. 4 qtrs | | | | 0.299* | |
| | | | | (0.160) | |
| Ext. G Debt to GDP | | | | | -0.009 |
| | | | | | (0.131) |
| Observations | 24 | 24 | 24 | 24 | 17 |
| R-squared | 0.102 | 0.102 | 0.107 | 0.269 | 0.052 |

[*] Note: This table presents the OLS coefficients of regressing the estimated two semester cumulative response of government bond yields to government innovations against income inequality (from OECD database), financial openness (from Lane and Milesi-Ferretti (2007)), inflation targeting dummy (Carare and Stone (2003)), foreign government debt to GDP (from IMF-World Bank QPSD data), and the fiscal multiplier. The fiscal multiplier is estimated analogously to the IRRF from the Blanchard and Perotti (2002)'s approach. We drop the Greece and the U.S. from the sample. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Table 11
IRRF not var. adj. and Household Debt Auerbach and Gorodnichenko (2013)'s shocks

| VARIABLES | (1) IRRF | (2) IRRF | (3) IRRF | (4) IRRF |
|---|---|---|---|---|
| HH debt to income | -0.092 | -0.131 | -0.097 | -0.200 |
| | (0.133) | (0.155) | (0.140) | (0.143) |
| Financial Openness | | 0.097 | | |
| | | (0.097) | | |
| Inflation Targeting | | | -0.144 | |
| | | | (0.323) | |
| Fiscal Mult. 4 qtrs | | | | 0.389** |
| | | | | (0.175) |
| Observations | 24 | 24 | 24 | 24 |
| R-squared | 0.016 | 0.032 | 0.026 | 0.291 |

* Note: This table presents the OLS coefficients of regressing the estimated two semester cumulative response of government bond yields to government innovations against household debt (from OECD database), financial openness (from Lane and Milesi-Ferretti (2007)), inflation targeting dummy (Carare and Stone (2003)), and the fiscal multiplier. The fiscal multiplier is estimated analogously to the IRRF from the Blanchard and Perotti (2002)'s approach. We drop Greece and the U.S from the sample. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

# Appendix B: Credit Constraints, IRRF, and Inequality

Here we study a version of our model with credit constrained households. What if instead of the minimum consumption level, which makes non-rich agents saving-constrained, the friction is a borrowing constraint? In particular, suppose that the constraint $c_0 \geq \underline{c}$ is replaced with

$$R\left(y^i + G - c_0\right) \geq \underline{b},$$

which says that the agents can at $t = 0$ promise to pay at most $-\underline{b} \geq 0$ at $t = 1$. If this constraint binds only for the non-rich, we have

$$c_0^p = y^p + G - \frac{1}{R}\underline{b}$$
$$c_0^r = \frac{1}{2}(1 - y^r)\, G + \frac{1}{2} y^r \left(1 + \frac{1}{R}\right),$$

and then market clearing ($\pi c_0^p + (1-\pi)\, c_0^r = 1$), $\Pi + w\ell^* = A\ell^{*\alpha} = 1$, and the definition of $y^i$ imply

$$1 = \pi y^p + \pi G - \pi \frac{1}{R}\underline{b} + (1-\pi)\frac{1}{2}(1-y^r)\, G + (1-\pi)\frac{1}{2}y^r \left(1 + \frac{1}{R}\right)$$

$$\Longrightarrow$$

$$\frac{1}{R} = \frac{\pi y^p + \pi G + (1-\pi)\frac{1}{2}(1-y^r)\, G - 1 + (1-\pi)\frac{1}{2}y^r}{\pi\underline{b} - (1-\pi)\frac{1}{2}y^r}$$

$$\frac{1}{R} = \frac{\pi y^p + \left[\pi + (1-\pi)\frac{1}{2} - (1-\pi)\frac{1}{2}y^r\right] G - 1 + (1-\pi)\frac{1}{2}y^r}{\pi\underline{b} - (1-\pi)\frac{1}{2}y^r}.$$

Thus,

$$\begin{aligned}\frac{\partial^2 (1/R)}{\partial G} &= \frac{\pi + (1-\pi)\frac{1}{2} - (1-\pi)\frac{1}{2}y^r}{\pi\underline{b} - (1-\pi)\frac{1}{2}y^r} \\ &= \frac{-\pi - (1-\pi)\frac{1}{2} + (1-\pi)\frac{1}{2}y^r}{-\pi\underline{b} + (1-\pi)\frac{1}{2}y^r} \\ &= \frac{-\pi - (1-\pi)\frac{1}{2} + (1-\pi)\frac{1}{2}\left(\frac{\Pi}{2(1-\pi)} + w\ell^*\right)}{-\pi\underline{b} + (1-\pi)\frac{1}{2}\left(\frac{\Pi}{2(1-\pi)} + w\ell^*\right)} \\ &= \frac{\frac{\Pi}{2} + (1-\pi)\, w\ell^* - 1 - \pi}{2\pi\,(-\underline{b}) + (1-\pi)\, w\ell^* + \frac{\Pi}{2}} < 0,\end{aligned}$$

where the inequality follows from $(1-\pi)y^r < 1$. And, if $-\underline{b}$ is small,

$$\frac{\partial^2 (1/R)}{\partial G \partial \pi} < 0$$

$$\Longrightarrow$$

$$\frac{\partial^2 R}{\partial G \partial \pi} > 0.$$

Therefore, with credit constrained households, the interest rate rises in response to a $G$ shock and the effect is amplified by inequality, contrary to our empirical findings.