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FISCAL CONSOLIDATION PROGRAMS AND INCOME INEQUALITY*

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We document a strong empirical relationship between higher income inequality and stronger recessive impacts of fiscal consolidation episodes across time and space. To explain this finding, we develop a life-cycle economy with uninsurable income risk. We calibrate our model to match key characteristics of several European economies, including inequality and fiscal structures, and study the effects of fiscal consolidation programs. In our model, higher income risk induces precautionary savings behavior, which decreases the proportion of credit-constrained agents in the economy. These agents have less elastic labor supply responses to fiscal consolidations, which explain the correlation with inequality in the data.

1. INTRODUCTION

The 2008 financial crisis led several European economies to adopt counter-cyclical fiscal policy, often financed by debt. Government deficits exceeded 10% in many countries, and this created an urgency for fiscal consolidation policies as soon as times returned to normal. In response, some governments designed plans to reduce their debt through austerity, tax increases, or, more commonly, a combination of the two—see Blanchard and Leigh (2013) and Alesina et al. (2015a). However, the process of fiscal consolidation across European countries raised a number of important questions about the effects on the economy. For example, is debt consolidation ultimately contractionary or expansionary? How large are the effects and do they depend on the state of the economy? How does the impact of consolidation through

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austerity differ from the impact of consolidation through taxation? This article contributes to this literature, both empirically and theoretically, by presenting evidence for a dimension that helps explain the heterogeneous responses to fiscal consolidations observed across countries: income inequality and, in particular, *the role of uninsurable income risk*.[2]

We begin by documenting a strong positive empirical relationship between higher income inequality and stronger recessive impacts of fiscal consolidation programs across time and place. We do this by using data and methods from three recent, state-of-the-art, empirical papers, which cover various countries and time periods and make use of different empirical approaches: (i) Blanchard and Leigh (2013), (ii) Alesina et al. (2015a), and (iii) Ilzetzki et al. (2013).[3]

Next we study the effects of fiscal consolidation programs, financed through both austerity and taxation, in a neoclassical macro-model with heterogeneous agents and incomplete insurance markets. We show that such a model is well-suited to explain the relationship between income inequality and the recessive effects of fiscal consolidation programs. The mechanism we propose works through idiosyncratic income risk. In economies with lower income risk, there are more credit-constrained households and households with low wealth levels, due to less precautionary saving. Importantly, these credit-constrained households have less elastic labor supply responses to increases in taxes and decreases in government expenditures.

Our empirical analysis begins with a replication of the recent studies by Blanchard and Leigh (2013, 2014). These studies find that the International Monetary Fund (IMF) underestimated the impacts of fiscal consolidation across European countries, with stronger consolidation causing larger GDP forecast errors. In Blanchard and Leigh (2014), the authors find no other significant explanatory factors, such as pre-crisis debt levels[4] or budget deficits, banking conditions, or a country's external position, among others, can help explain the forecast errors. In Subsection 3.1, we reproduce the exercise conducted by Blanchard and Leigh (2013), now augmented with different metrics of income inequality. We find that during the 2010 and 2011 consolidation in Europe, the forecast errors are larger for countries with higher income inequality, implying that inequality amplified the recessive impacts of fiscal consolidation. A one standard deviation increase in income inequality, measured as Y_{10}/Y_{90}[5] leads the IMF to underestimate the fiscal multiplier in a country by 66%.

For a second independent analysis, we use the Alesina et al. (2015a) fiscal consolidation episodes data set with data from 12 European countries over the period 2007–2013. Alesina et al. (2015a) expand the exogenous fiscal consolidation episodes data set, known as IMF shocks, from Devries et al. (2011) who use Romer and Romer (2010) narrative approach to identify exogenous shifts in fiscal policy. Again we document the same strong amplifying effect of income inequality on the recessive impacts of fiscal consolidation. A one-standard deviation increase in inequality, measured as Y_{25}/Y_{75}, increases the fiscal multiplier by 240%.

Our third empirical analysis (for brevity included in the Appendix) replicates the paper by Ilzetzki et al. (2013). These authors use time series data from 44 countries (both rich and poor) and a SVAR approach to study the impacts of different country characteristics on fiscal multipliers. We find that countries with higher income inequality experience significantly stronger declines in output following decreases in government consumption.

To explain these empirical findings, we develop an overlapping generations economy with heterogeneous agents, exogenous credit constraints, and uninsurable idiosyncratic risk. We

[2] In this article, we focus on two main drivers of income inequality: permanent and transitory income shocks. Recently, there has been much attention paid to job polarization and job displacement by automation as a source of increases in income inequality over time. However, given the cross-sectional nature of our study, we abstract from such mechanism.

[3] Although the first two papers study fiscal consolidation programs in Europe, Ilzetzki et al. (2013) study government spending multipliers using a greater number of countries. We include this study for robustness and completeness.

[4] In Section 8.1, we show that, in line with our proposed mechanism, household debt matters if an interaction term between debt and the planned fiscal consolidation is included in the regression.

[5] Ratio of top 10% income share over bottom 10% income share.

calibrate the model to match data from a number of European countries along dimensions such as the distribution of income and wealth, taxes, social security, and debt level. Then we study how these economies respond to gradually reducing government debt, either by cutting government spending or by increasing labor income taxes.

Output falls when debt reduction is financed through either a decrease in government spending or increased labor income taxes. In both cases, this is caused by a fall in labor supply. In the case of reduced government spending, the transmission mechanism works through a future income effect. As government debt is paid down, the capital stock and thus the marginal product of labor (wages) rise, and thus expected lifetime income increases. This will lead agents to enjoy more leisure and decrease their labor supply today,[6] and output to fall in the short run, despite the long-run effects of consolidation on output being positive. Credit-constrained agents and agents with low wealth levels do, however, have a lower marginal propensity to consume goods and leisure out of future income (for constrained agents, the marginal propensity to consume out of future income is zero.)[7] Constrained agents have a lower intertemporal elasticity of substitution and do not consider changes to their lifetime budget, only changes to their budget in the current time period.[8]

In the case of consolidation through increased labor income taxes, there will be a negative substitution effect on labor supply today, in addition to an income effect that could be positive or negative depending on whether the future taxes or higher future wages dominate. Because wages are rising in the future, unconstrained agents would prefer to work relatively less today. For constrained agents, who do not consider their life-time budget but only their budget today, the tax would cause a drop in available income in the short run (no future wage increases to be considered), possibly leading to a labor supply increase. It turns out that all agents decrease their labor supply, but the response is weaker for constrained and low-wealth agents.

When higher income inequality reflects higher uninsurable income risk, there exists a negative relationship between income inequality and the number of credit-constrained agents. Greater risk leads to increased precautionary savings behavior, thereby decreasing the share of agents with liquidity constraints and low wealth levels. Since unconstrained agents have a higher intertemporal elasticity of substitution of labor and thus a more elastic labor supply response to both tax-based and austerity-based consolidation, labor supply and output will respond more strongly in economies with higher inequality.

Through simulations in a benchmark economy, initially calibrated to Germany, we show that varying the level of idiosyncratic income risk strongly affects the fraction of credit-constrained agents in the economy and the fiscal multiplier, both for consolidation through taxation and austerity. If we instead change inequality by changing the variance of initial conditions, prior to entering the labor market (permanent ability and the age profile of wages in the model), there is very little effect on the fraction of credit-constrained agents or on the fiscal multiplier.

In a multicountry exercise, we calibrate our model to match a wide range of data and country-specific policies from 13 European economies, and find that our simulations reproduce the anticipated cross-country correlation between income inequality and fiscal multipliers. Moreover, we show that in our model, countries with higher idiosyncratic uninsurable labor income risk have a smaller percentage of constrained agents and have larger multipliers, confirming our analysis and mechanism for the benchmark model calibrated to Germany.

One should note that our mechanism relies on the premise that differences in income inequality across countries are, at least to some extent, explained by differences in the amount of uninsurable risk that agents in different countries are exposed to. The stronger the

[6] A recent paper by Cesarini et al. (2017) provides empirical evidence on how positive wealth shocks lead to a fall in labor supply.

[7] The fact that constrained agents also very slightly change their labor supply in our model simulations is due to general equilibrium effects (price changes) today.

[8] Domeij and Floden (2006) provide empirical evidence of the intertemporal elasticity of substitution of labor being decreasing in wealth, using U.S. micro-data.

correlation between risk and inequality in reality, the stronger the proposed mechanism will be.

Although we do not have direct evidence on cross-country differences in income risk, we do perform a number of empirical exercises whose results are consistent with our model predictions, under the premise above.[9]

Furthermore, quantitatively the mechanism depends crucially on general equilibrium effects, see Subsection 7.2. This is especially true in the case of austerity-based consolidation, where the effect comes through future changes in prices. If debt consolidation does not lead to the crowding in of productive capital (as would be the case in a small open economy with an interest rate set in the international market), the size of the fiscal multipliers becomes essentially zero, although the correlation with wage inequality is still there.[10] In the case of tax-based consolidation, the fiscal multiplier becomes larger in partial equilibrium and the cross-country correlation with income risk remains high.[11]

We perform two empirical exercises to test the validity of the mechanism described above. First, in our calibrated model, higher levels of household debt are associated with a higher number of credit-constrained households. This implies that countries with higher levels of household debt should have experienced less recessive impacts of fiscal consolidation programs. We show that this relationship exists in the data, by again performing a similar exercise to Blanchard and Leigh (2013).

Second, the mechanism we propose is dependent on poor agents having a lower labor supply elasticity to fiscal consolidation shocks than relatively wealthy agents. We use the Panel Study of Income Dynamics (PSID) data together with fiscal consolidation shocks identified by Alesina et al. (2015a) and analyze how labor supply responds to fiscal consolidation shocks and how this depends on household wealth. We find that poor agents have lower labor supply responses to fiscal shocks when government debt is changing, precisely as our model predicts.

In Section 9, we conduct a final validity test of the mechanism by using our model. In the empirical analysis, we make the case that the IMF forecasts did not properly take income inequality into account. In this section, we show that using data from our model, obtained by simulating the observed fiscal consolidation shocks in the data, we get similar results to Blanchard and Leigh (2013) when we shut down all labor income risk in our model. The difference between the output drop that our calibrated model predicts both including and excluding income risk (which is our proxy for the forecast error under the assumption that the forecast was misspecified by excluding inequality), is explained by the size of the fiscal shock and its interaction with the same income inequality metrics as in our replication of the Blanchard and Leigh (2013) experiment (found in Subsection 3.1). The resulting pattern of regression statistics is strikingly similar to Blanchard and Leigh (2013).

The remainder of the article is organized as follows: We begin by discussing some of the recent relevant literature in Section 2. In Section 3, we assess the empirical relationship between income inequality and the fiscal multipliers associated with consolidation programs. In Section 4, we describe the overlapping generations model, define the competitive equilibrium, and explain the fiscal consolidation experiments. Section 5 describes the calibration of the model. In Section 6, we inspect the transmission mechanism, followed by the cross-country analysis in Section 7. In Section 8, we empirically validate the mechanism and in Section 9 we replicate the Blanchard and Leigh (2014) exercise with model data. Section 10 concludes.

[9] We also perform a number of robustness exercises with regards to the share of income inequality that is explained by permanent differences across individuals and by uninsurable risk and our results still hold up as long as a significant fraction is explained by risk.

[10] There is still a very small fiscal multiplier due to agents receiving a greater lump-sum transfer at the end of the consolidation period and forever after—50 years in our exercises—resulting from a lower debt service.

[11] In Subsection 7.2, we show that in partial equilibrium the average multiplier in the tax-based consolidation increases in absolute value from 1.27 in the benchmark to 2.71, even if future wages are no longer rising. The labor supply of unconstrained agents responds more strongly. They have some savings or borrowing capacity from before the unexpected tax increase, which they can use to smooth consumption and they reduce labor supply more than constrained agents who have to work hard today to avoid a large drop in consumption.

2. RELATED LITERATURE

This article relates to several branches of the literature in the area of fiscal policy and inequality. First, to a burgeoning literature that focuses on the impacts that consolidation programs have on output. Papers such as Guajardo et al. (2014) and Yang et al. (2015) find a negative impact on output stemming from fiscal consolidation shocks. Blanchard and Leigh (2013) and Blanchard and Leigh (2014) find that during the sovereign debt crisis in Europe, the implemented fiscal consolidation programs had a recessive effect on output and show that this effect is underestimated by the IMF. The conclusions in Alesina et al. (2015b) support previous studies, emphasizing that tax-based consolidations produce deeper and longer recessions than spending based ones.

Closely related to this literature is the one assessing the welfare costs of consolidation programs. Röhrs and Winter (2017) suggest that even though there are long-term welfare benefits of fiscal consolidation in the United States, they do not outweigh the welfare costs of the transition to the new steady state. The authors also find evidence that wealth inequality is a major driver of welfare effects. Romei (2015) addresses the issue of the optimal speed and composition of a fiscal consolidation, concluding that it should be done quickly and by cutting public expenditure.

Second, our article relates to the recent literature assessing the determinants of the fiscal multipliers. Heathcote (2005) studies the effects of changes in the timing of income taxes and finds that tax cuts can have large real effects and that the magnitude of the effect depends crucially on the degree of market incompleteness. Hagedorn et al. (2016), in a New Keynesian model, present further evidence of the relevance of market incompleteness in determining the size of fiscal multipliers. Ferriere and Navarro (2016) provide empirical evidence showing that in the postwar United States, fiscal expansions are only expansionary when financed by increases in tax progressivity.

The closest article to ours is Brinca et al. (2016b), where the authors show that the interaction between the distribution of wealth and the labor supply response to a fiscal shock can be a determinant of the size of the fiscal multiplier. They study a one-period, balanced budget, government expenditure expansion, financed by a lump-sum tax. Using a standard incomplete markets model, the authors show that observed cross-country differences in the wealth distribution can lead to different shares of credit-constrained agents and lead to different multipliers.[12] In this article, we make two important contributions to this framework: first, we focus on more empirically plausible fiscal policies that have been at the center of the recent policy debate, namely debt consolidation events; second, we focus on another source of cross-country heterogeneity that can also affect the share of credit-constrained agents in a given economy, namely income risk. It should be noted in this context that in the data the cross-country correlation between income and wealth inequality is relatively weak.[13] In the context of our model, higher income dispersion leads to fewer credit-constrained agents, due to precautionary savings behavior. Higher wealth inequality, induced by heterogeneity in discount factors, in contrast, typically leads to more credit-constrained agents.

Finally, this article is closely related to the literature that establishes the importance of the distribution of wealth (credit constraints in particular) for the response of labor and consumption to income shocks. Domeij and Floden (2006) document both empirically, using data from the PSID, and quantitatively, using an incomplete markets macro model, that the intertemporal elasticity of substitution of labor is increasing in wealth. This finding is at the heart of the mechanism in our article. Further evidence of heterogeneous responses is provided by Anderson et al. (2016) who find that, in the context of the U.S. economy, individuals respond

[12] Data on credit-constrained agents are not available for many countries and both the current paper and Brinca et al. (2016b) have to vary the model parameters governing some measures of wealth or income inequality and observe what this does to the fraction of credit-constrained agents in the economy.

[13] For the 26 European countries used in Ilzetzki et al. (2013), we find that the correlation between the wealth and income Gini is close to 0.

differently to unanticipated fiscal shocks depending on age, income level, and education. The behavior of the highest earners, in particular, is consistent with Ricardian equivalence, while poor households show evidence of non-Ricardian behavior. Krueger et al. (2016) assess how wealth, income, and preference heterogeneity across households amplify aggregate shocks, and conclude that the income and wealth distribution play a crucial role for the response of consumption. Pham-Dao (2019) focuses on the role of social security and means-tested public transfers as a determinant of precautionary savings behavior, which is a key mechanism in our analysis. In addition to income risk, public policy may lead to cross-country differences in the response to fiscal consolidations.[14]

3. EMPIRICAL ANALYSIS

In this section, we document a strong empirical relationship between income inequality and the fiscal multiplier resulting from fiscal consolidation programs. We do this by augmenting recent exercises by Blanchard and Leigh (2013) and Alesina et al. (2015a), who study the impact of recent fiscal consolidation shocks, to account for the effects of income inequality in the propagation of fiscal consolidation shocks. In the Appendix, we show that income inequality is an important dimension when studying fiscal shocks in general and not only consolidations.

3.1. *GDP Forecast Errors and Fiscal Consolidation Forecasts.* Blanchard and Leigh (2013) propose a standard rational expectations model specification to investigate the cross-country relationship between growth forecast errors and planned fiscal consolidations after the crisis. The approach consists of regressing forecast errors for real GDP growth on forecasts of fiscal consolidations made in the beginning of 2010. The specification proposed by Blanchard and Leigh is the following:

$$\Delta Y_{i,t:t+1} - \hat{E}\{\Delta Y_{i,t:t+1}|\Omega_t\} = \alpha + \beta \hat{E}\{F_{i,t:t+1|t}|\Omega_t\} + \epsilon_{i,t:t+1}, \tag{1}$$

where α is a constant, $\Delta Y_{i,t:t+1}$ is the cumulative year-to-year GDP growth rate in economy i from period t to $t+1$ (years 2010 and 2011, respectively), and the forecast error is measured as $\Delta Y_{i,t:t+1} - \hat{E}\{\Delta Y_{i,t:t+1}|\Omega_t\}$, with $\hat{E}$ being the forecast conditioned on the information set Ω at time t. $\hat{E}\{F_{i,t:t+1|t}|\Omega_t\}$ denotes the planned cumulative change in the general government structural fiscal balance in percentage of potential GDP, and is used as a measure of discretionary fiscal policy.

Under the null hypothesis that the IMF's forecasts regarding the impacts of fiscal consolidation were accurate, β should be zero. What Blanchard and Leigh (2013) find is that β is not only statistically different from zero, but negative and around 1. This means that the IMF severely underestimated the recessive impacts of austerity, implying that for every additional percentage point of fiscal consolidation, output was about 1% lower than what was forecast.[15]

Blanchard and Leigh (2013) then investigate what other factors could explain the forecast errors. The authors test for initial level of financial stress, initial level of external imbalances, trade-weighted forecasts of trading partners' fiscal consolidation forecasts, the initial level of household debt,[16] the IMF's Early Warning exercise vulnerability ratings computed in early 2010, and other variables. The results are robust and no control is significant. Two conclusions are drawn from this. First, that none of the variables examined are correlated with both the forecast error and planned fiscal consolidation and thus the underestimation of the recessive impacts of consolidation are not related with these different dimensions. Second, since none

[14] An interesting extension, that is beyond the scope of this article, would be to investigate the impact of welfare state policies on fiscal multipliers.

[15] Blanchard and Leigh (2013) also investigate whether this result could have been driven by the fact that planned fiscal consolidations were different from actual ones. The authors show that this was not the case, as planned and actual consolidations have a correlation close to one.

[16] In Section 8, we show that household debt matters if interacted with the planned fiscal consolidation.

TABLE 1
BLANCHARD AND LEIGH (2013) REGRESSIONS AUGMENTED WITH MEASURES OF INCOME INEQUALITY

Coefficients	(1) Blanchard–Leigh	(2) Y25/Y75	(3) Y20/Y80	(4) Y10/Y90	(5) Y5/Y95	(6) Y2/Y98	(7) Income Gini
β	−1.095***	−0.841***	−0.806***	−0.697**	−0.759***	−0.750***	−1.267***
	(0.255)	(0.227)	(0.234)	(0.252)	(0.240)	(0.238)	(0.275)
γ		−0.194	−0.144	−0.065	0.008	0.018	0.273**
		(0.385)	(0.291)	(0.120)	(0.036)	(0.032)	(0.121)
ι		−0.251	−0.238	−0.154***	−0.071***	−0.066***	−0.085
		(0.208)	(0.153)	(0.054)	(0.021)	(0.019)	(0.084)
Constant	0.775*	2.150	2.041	1.812	0.805	0.558	−9.344**
	(0.383)	(2.632)	(2.422)	(1.758)	(0.928)	(0.597)	(4.463)
Observations	26	26	26	26	26	26	26
R^2	0.496	0.545	0.559	0.612	0.600	0.610	0.624

NOTES: The table displays the results from augmenting the regression in Blanchard and Leigh (2013) with different measures of income inequality and an interaction term between income inequality and planned fiscal consolidation. Y25/Y75, Y20/Y80, Y10/Y90, Y5/Y95, and Y2/Y98 represent the share of income of the top 25%, 20%, 10%, 5%, and 2% divided by the share of the bottom 25%, 20%, 10%, 5%, and 2%. Robust standard errors in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, and $^{*}p < 0.1$.

of these factors are statistically significant, we can infer that none of them significantly affected the forecast errors of the IMF.

We expand Equation (1) to account for several different metrics of income inequality.[17] Using the European Union Statistics on Income and Living Conditions (EU-SILC) data set, we construct various measures of income inequality for the same 26 European economies used by Blanchard and Leigh (2013).[18]

Moreover, to test whether inequality helps to explain the impact of fiscal consolidation, we include in the regression an interaction between the planned fiscal consolidation and inequality. To provide better intuition, we reparametrize the specification and demean the inequality measures in the interaction term. Therefore, we estimate the following equation:

$$\begin{aligned} \Delta Y_{i,t:t+1} - \hat{E}\{\Delta Y_{i,t:t+1}|\Omega_t\} &= \alpha + \beta \hat{E}\{F_{i,t:t+1|t}|\Omega_t\} + \gamma I_{i,t-1} \\ &\quad + \iota((\hat{E}\{F_{i,t:t+1|t}|\Omega_t\})(I_{i,t-1} - \mu_I)) + \epsilon_{i,t:t+1}, \end{aligned} \tag{2}$$

where $I_{i,t-1}$ is the inequality measure for country i and μ represents the mean of I. We use lagged inequality to guarantee that it is not influenced by GDP growth rate or by the fiscal consolidation measures. The results are presented in Table 1. When the demeaned inequality measures are included, the β coefficients have a convenient interpretation as how much the effects of fiscal consolidation were underestimated for a country with inequality equal to the sample mean. The ι coefficients tell us by how much more (relative to the β coefficients) the IMF underestimated the fiscal consolidation effects for a country with inequality 1 percentage point above the sample mean.

First, relative to the benchmark case of Blanchard and Leigh (2013), we see that even though the consolidation variable is still statistically significant, the coefficient point estimates are now smaller in absolute value. This tells us that including income inequality and its

[17] The shares of income of top 25%, 20%, 10%, 5%, and 2% over the share of the bottom 25%, 20%, 10%, 5%, and 2%, respectively, and the income Gini coefficient. In Table A6, we test with other measures of income inequality such as the share of income of the bottom 25%, 20%, 10%, and 5%. Results are robust to these alternative measures. Here we use gross income but in Table A7 we provide evidence that the result is robust to using disposable income instead.

[18] The 26 economies used by Blanchard and Leigh were Austria, Belgium, Bulgaria, Cyprus, Czech Republic, Germany, Denmark, Finland, France, Greece, Hungary, Ireland, Iceland, Italy, Malta, Netherlands, Norway, Poland, Portugal, Romania, Slovak Republic, Slovenia, Spain, Sweden, Switzerland, and the United Kingdom.

interaction with planned consolidation reduces the impacts of the size of fiscal consolidation in itself.

Second, note that an increase of 1% above the mean of income inequality amplifies the forecast error of the effects of fiscal consolidation by ι. This means that if the forecasters had taken income inequality into account, the effects of fiscal consolidation would have been more accurately anticipated.

The results are not only statistically significant and robust, but also economically meaningful. For example, an increase in one standard deviation of the income share of agents in the top 10% of the income distribution over the bottom 10% leads to an underestimation of the fiscal multiplier of 66%, for a country with an average consolidation.[19,20]

3.2. *IMF Shocks.* In this subsection, we show that the link between income inequality and the output response to fiscal consolidations is not exclusive to the years of 2010 and 2011. We use the Alesina et al. (2015a) annual data set on fiscal consolidation episodes in 12 European economies[21] between 1978 and 2013. The authors expand the exogenous fiscal consolidation episodes data set in Devries et al. (2011), known as IMF shocks, which is constructed using the Romer and Romer (2010) narrative approach to identify fiscal consolidations solely driven by the need to reduce deficits. The use of the narrative approach makes it possible to filter out all policy actions driven by the economic cycle and guarantees exogeneity of the shifts in fiscal policy.

Alesina et al. (2015a) expand the Devries et al. (2011) data set, but use the methodological innovation proposed by Alesina et al. (2015b), who notice that a fiscal adjustment is not an isolated change in expenditure or taxes, it is a multiyear plan, in which some policies are known in advance and others are implemented unexpectedly. Ignoring the connection between the unanticipated and announced consolidation measures can lead to biased results.

In the Alesina et al. (2015a) data set, fiscal consolidations are measured as expected revenue effects of changes in the tax code and as deviations of expenditure relative to the expected level of expenditure absent the policy changes. The fiscal consolidation episodes are assumed to be fully credible, and announcements that were not implemented are dropped from the database.

Once again, we use total income inequality data from the EU-SILC data set and construct the same measures of income inequality as in Subsection 3.1. The EU-SILC data go from 2007 to 2015 for all the 12 European economies in the Alesina et al. (2015a) data set. The equation that we estimate is the following:

$$(3)\quad \Delta Y_{i,t} = \alpha + \beta_1 e^u_{i,t} + \beta_2 e^a_{i,t} + \gamma I_{i,t-1} + \iota_1 e^u_{i,t}(I_{i,t-1} - \mu_I) + \iota_2 e^a_{i,t}(I_{i,t-1} - \mu_I) + \delta_i + \omega_t + \epsilon_{i,t},$$

where ΔY_{it} is the GDP growth rate in economy i in year t, e^u_{it} is the unanticipated consolidation shock while e^a_{it} is the announced shock. I_{it-1} is the inequality measure in year $t-1$ and μ represents the sample mean of I. We consider the lagged value of inequality to guarantee that inequality is not affected by current changes in output and current fiscal consolidation. We reparameterize the interaction terms by demeaning the inequality measures so that β_1 and β_2 have the more convenient interpretation of how much, for a country with average inequality, an increase in fiscal consolidation of 1% affects output growth for a country with average inequality. Moreover, ι_1 and ι_2 also have the more convenient interpretation of by how much more (relative to a country with average inequality) fiscal consolidation affects the

[19] Note that even though this is a statement about the IMF's forecast errors, if we use output alone as the dependent variable, we still find similar results, showing that higher income inequality is associated with a higher impact of fiscal consolidation, see Table A8.

[20] The results are robust to both the inclusion of a lagged term of the forecast error and to winsorizing observations at the 95th percentile, see Tables A9 and A10.

[21] Austria, Belgium, Germany, Denmark, Spain, Finland, France, United Kingdom, Ireland, Italy, Portugal, and Sweden.

TABLE 2
REGRESSIONS ON DATA FROM ALESINA ET AL. (2015A)

Coefficients	(1) Benchmark	(2) Y25/Y75	(3) Y20/Y80	(4) Y10/Y90	(5) Y5/Y95	(6) Y2/Y98	(7) Income Gini
β_1	−0.003	0.006	0.004	−0.004	−0.004	−0.004	0.011
	(0.005)	(0.007)	(0.007)	(0.006)	(0.006)	(0.007)	(0.007)
β_2	−0.002	−0.003	−0.002	−0.000	−0.002	0.001	−0.001
	(0.005)	(0.007)	(0.007)	(0.007)	(0.006)	(0.006)	(0.007)
γ		−2.294**	−1.308*	−0.024	0.036	0.009	−1.100***
		(1.001)	(0.756)	(0.344)	(0.135)	(0.049)	(0.380)
ι_1		−1.363**	−0.882*	0.103	0.069	−0.005	−0.501**
		(0.590)	(0.501)	(0.232)	(0.077)	(0.030)	(0.191)
ι_2		−0.357	−0.213	−0.094	−0.017	0.022	−0.112
		(0.633)	(0.510)	(0.245)	(0.091)	(0.026)	(0.173)
Constant	0.014***	0.171**	0.123*	0.018	0.005	0.012	0.434***
	(0.005)	(0.069)	(0.063)	(0.050)	(0.034)	(0.014)	(0.145)
Observations	84	84	84	84	84	84	84
R^2	0.008	0.132	0.086	0.012	0.030	0.021	0.179
Number of countries	12	12	12	12	12	12	12

NOTES: The table displays the results from estimating the regression in Equation (3) on data from Alesina et al. (2015a) and measures of income inequality from the EU-SILC. Y25/Y75, Y20/Y80, Y10/Y90, Y5/Y95, and Y2/Y98 represent the share of income of the top 25%, 20%, 10%, 5%, and 2% divided by the share of the bottom 25%, 20%, 10%, 5%, and 2%. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Standard errors in parentheses.

GDP growth rate for a country with inequality 1 percentage point above the sample mean. δ_i and ω_t are country and year fixed effects.

The results are presented in Table 2. Note that from the two interaction terms, only the interaction with unanticipated IMF shocks is statistically significant. This tells us that, for an unanticipated fiscal consolidation, an increase in inequality by 1 percentage point is going to amplify the recessive impacts of fiscal consolidation (the fiscal multiplier) by ι_1.

Once again, the results are not only robust and statistically significant, but also economically meaningful. An increase of one standard deviation in the share of the income of the top 25% over the share of the bottom 25% leads to an increase in the multiplier of an unanticipated shocks of 240%, for a country with an average unanticipated consolidation.[22]

The empirical findings in Section 3 together suggest that income inequality is a relevant dimension to take into account when studying the effects of fiscal policy.[23] In particular, they suggest that higher inequality amplifies the recessive impacts of fiscal consolidation and decreases in government expenditures. To understand the mechanism through which income inequality may play such role, we build a structural model that is introduced in the next section.

4. MODEL

In this section, we describe the model we will use to study the effects of a fiscal consolidation in different countries. Our model is a relatively standard life-cycle economy with heterogeneous agents and incomplete markets, that is, a life cycle extension of Aiyagari (1994). It is similar to the model in Brinca et al. (2016b), except that we have introduced a bequest motive to get a more realistic distribution of wealth over the life cycle.

4.1. *Technology.* We use the standard assumption of a representative firm, producing output with a Cobb–Douglas production function, see the Appendix for a detailed description.

[22] Once again, the results are robust to both the inclusion of the lagged dependent variable and to winsorizing. See Tables A11 and A12.

[23] Further evidence, with one other independent data set and identification strategy, in the Appendix support the findings in this section.

4.2. *Demographics.* The economy is populated by J overlapping generations of finitely lived households.[24] All households start life at age 20 and enter retirement at age 65. Let j denote the household's age. Retired households face an age-dependent probability of dying, $\pi(j)$ and die for certain at age 100.[25] A model period is one year, so there are a total of 40 model periods of active work life. We assume that the size of the population is fixed (there is no population growth). We normalize the size of each new cohort to 1. Using $\omega(j) = 1 - \pi(j)$ to denote the age-dependent survival probability, by the law of large numbers the mass of retired agents of age $j \geq 65$ still alive at any given period is equal to $\Omega_j = \prod_{q=65}^{q=J-1} \omega(q)$.

In addition to age differences, households are heterogeneous with respect to asset holdings, idiosyncratic productivity, and their subjective discount factor, which for each household is constant over time but takes one out of the three values $\beta \in \{\beta_1, \beta_2, \beta_3\}$. We follow Krusell and Smith (1998) in using heterogeneity in discount factors to allow the model to match the empirical wealth distribution. Such an approach is supported by empirical evidence in Mischel et al. (1972) and more recently by Epper et al. (2020), who find that patience is as important as education in predicting a person's position in the wealth distribution. The discount factors are distributed uniformly across agents in each cohort. Finally, they also differ in terms of a permanent ability component, that is, they have a starting level of productivity that is realized at birth. Every period of active work-life they decide how many hours to work, n, how much to consume, c, and how much to save, k. Retired households make no labor supply decisions but receive a social security payment, Ψ_t.

There are no annuity markets, so that a fraction of households leave unintended bequests, which are redistributed in a lump-sum manner between the households that are currently alive. We use Γ to denote the per-household bequest. Retired households' utility is increasing in the bequest they leave when they die. This helps us calibratem the asset holdings of old households.

4.3. *Labor Income.* The wage of an individual depends on his/her own characteristics: age, j, permanent ability, $a \sim N(0, \sigma_a^2)$, and idiosyncratic productivity shock, u, which follows an AR(1) process:

$$u_{t+1} = \rho u_t + \epsilon_{t+1}, \quad \epsilon \sim N(0, \sigma_\epsilon^2). \tag{4}$$

These characteristics will dictate the number of efficient units of labor the household is endowed with. Individual wages will also depend on the wage per efficiency unit of labor w. Thus, individual i's wage is given by

$$w_i(j, a, u) = w e^{\gamma_1 j + \gamma_2 j^2 + \gamma_3 j^3 + a + u}. \tag{5}$$

γ_1, γ_2, and γ_3 capture the age profile of wages.

4.4. *Preferences.* The momentary utility function of a household, $U(c, n)$, depends on consumption and work hours, $n \in (0, 1]$, in addition to a pure public good, G provided by the government, and takes the following form:

$$U(c, n) = \frac{c^{1-\sigma}}{1-\sigma} - \chi \frac{n^{1+\eta}}{1+\eta} + \log(G). \tag{6}$$

Retired households gain utility from the bequest they leave when they die:

$$D(k) = \varphi \log(k). \tag{7}$$

[24] Recent work by Peterman and Sager (2016) makes the case for having a life cycle dimension when studying the impacts of government debt.

[25] This means that $J = 81$.

4.5. *Government.* The government runs a balanced social security system where it taxes employees and the employer (the representative firm) at rates τ_{ss} and $\tilde{\tau}_{ss}$ and pays benefits, Ψ_t, to retirees. The government also taxes consumption and labor and capital income to finance the expenditures on pure public consumption goods, G_t, which enter separably in the utility function, interest payments on the national debt, rB_t, and a lump-sum redistribution, g_t. We assume that there is some outstanding government debt and that government debt-to-output ratio, $B_Y = B_t/Y_t$, does not change over time. Consumption and capital income are taxed at flat rates the τ_c and τ_k. To model the nonlinear labor income tax, we use the functional form proposed in Benabou (2002) and recently used in Heathcote et al. (2017) and Holter et al. (2019)

$$\tau_l(y) = 1 - \theta_0 y^{-\theta_1}, \tag{8}$$

where y denotes pre-tax (labor) income and $\tau_l(y)$ the average tax rate given a pre-tax income of y. The parameters θ_0 and θ_1 govern the level and the progressivity of the tax code, respectively.[26] Heathcote et al. (2017) argue that this function fits the U.S. data well.

In a steady state, the ratio of government revenues to output will remain constant. G_t, g_t, and Ψ_t must also remain proportional to output. Denoting the government's revenues from labor, capital, and consumption taxes by R_t and the government's revenues from social security taxes by R_t^{ss}, the government budget constraint in steady state takes the following form:

$$g\left(45 + \sum_{j\geq 65} \Omega_j\right) = R - G - rB, \tag{9}$$

$$\Psi\left(\sum_{j\geq 65} \Omega_j\right) = R^{ss}. \tag{10}$$

4.6. *Recursive Formulation of the Household Problem.* At any given time, a household is characterized by (k, β, a, u, j), where k is the household's savings, $\beta \in \beta_1, \beta_2, \beta_3$, is the time discount factor, a is permanent ability, u is the idiosyncratic productivity shock, and j is the age of the household. We can formulate the household's optimization problem over consumption, c, work hours, n, and future asset holdings, k', recursively as follows:

$$\begin{aligned}
V(k, \beta, a, u, j) &= \max_{c,k',n} \left[U(c, n) + \beta E_{u'}\left[V(k', \beta, a, u, j+1)\right]\right] \\
&\text{s.t.:} \\
c(1+\tau_c) + k' &= (k + \Gamma)(1 + r(1-\tau_k)) + g + Y^L \\
Y^L &= \frac{nw(j,a,u)}{1+\tilde{\tau}_{ss}}\left(1 - \tau_{ss} - \tau_l\left(\frac{nw(j,a,u)}{1+\tilde{\tau}_{ss}}\right)\right) \\
n \in [0,1], &\qquad k' \geq -b, \quad c > 0.
\end{aligned} \tag{11}$$

Here, Y^L is the household's labor income after social security taxes and labor income taxes. τ_{ss} and $\tilde{\tau}_{ss}$ are the social-security contributions paid by the employee and by the employer, respectively. The problem of a retired household, who has a probability $\pi(j)$ of dying and gains utility $D(k')$ from leaving a bequest, is

$$\begin{aligned}
V(k, \beta, j) &= \max_{c,k'} \left[U(c, n) + \beta(1-\pi(j))V(k', \beta, j+1) + \pi(j)D(k')\right] \\
&\text{s.t.:}
\end{aligned}$$

[26] A further discussion of the properties of this tax function is provided in the Appendix.

$$c(1+\tau_c)+k' = (k+\Gamma)(1+r(1-\tau_k))+g+\Psi,$$
$$k' \geq 0, \quad c > 0. \tag{12}$$

4.7. *Stationary Recursive Competitive Equilibrium.* In equilibrium, agents optimize at given prices, markets clear, budgets balance, and the cross-sectional distribution across household types is stationary. For the sake of brevity, the formal equilibrium definition is stated in the Appendix.

4.8. *Fiscal Experiment and Transition.* The fiscal experiments that we analyze in this article are 50 periods of reduction in government debt, B, either financed through a decrease in government spending, G, by 0.2% of benchmark GDP,[27] or an increase in the labor income tax τ_l, by 0.1% for all agents. The economy is initially in a steady state and the 50 periods of fiscal consolidation is unanticipated until it is announced.[28] After the 50 periods, either the government spending or the labor tax go back to the initial level. The lump-sum transfer, g, is fixed during the first 50 periods and then it is set to clear the government budget with the new level of debt and interest rate, and we assume that the economy takes an additional 50 periods to converge to the new steady state equilibrium, with a lower debt-to-GDP ratio.

To save space, the definition of a transition equilibrium after the fiscal experiment is stated in the Appendix. The key change compared to the steady state is that the dynamic-programming problem of households needs another state variable: time, t, capturing all the changes in policy and price variables relevant in this maximization problem. The numerical solution of the model necessitates guessing on paths for all the variables that will depend on time and then solving this maximization problem backward, after which the guess is updated; the method is similar to that used in Brinca et al. (2016b) and Krusell and Smith (1999).

4.9. *Definition of the Fiscal Multiplier in the Context of a Fiscal Consolidation Shock.* In the experiment with debt reduction financed by a reduction in G or an increase in R, we define fiscal multiplier as the net present value of the sum of all changes in output divided by the same period changes in the fiscal instrument (government consumption or tax revenues). Consequently, the impact multiplier is just the ratio of the change in output to the change in the fiscal instrument when the initial shock is realized (see the Appendix).

5. CALIBRATION

Our benchmark model is calibrated to match moments of the German economy. Germany is a natural choice as it is the largest economy in Europe. For the cross-country analysis in Section 7, calibration is performed using the same strategy and is presented in the Appendix. Certain parameters can be calibrated outside the model using direct empirical counterparts. Tables A21 and A23 list the parameters calibrated outside of the model. The remaining parameters, listed in Tables 4 (only Germany) and A22, are calibrated using a simulated method of moments (SMM) approach.

5.1. *Wages.* To estimate the life cycle profile of wages (see Equation (5)), we use data from the Luxembourg Income Study (LIS) and run the below regression for each country:

$$\ln(w_i) = \ln(w) + \gamma_1 j + \gamma_2 j^2 + \gamma_3 j^3 + \varepsilon_i, \tag{13}$$

where j is the age of individual i. The parameter for the variance of ability, σ_a, is assumed to be equal across countries and set equal to the average of σ_a for the European countries

[27] The total revenue available for debt repayment over the 50-year period is thus 10% of benchmark GDP.

[28] In Subsection 3.2, we find that unanticipated fiscal consolidations have a statistically significant negative effect on output, but anticipated consolidations do not.

TABLE 3
CALIBRATION FIT

Data Moment	Description	Source	Data Value	Model Value
$\bar{a}_{75-80}/\bar{a}$	Mean wealth age 75–80/mean wealth	LWS	1.51	1.51
K/Y	Capital-output ratio	PWT	3.013	3.013
$\mathrm{Var}(\ln w)$	Variance of log wages	LIS	0.354	0.354
$\bar{n}$	Fraction of hours worked	OECD	0.189	0.189
Q_{25}, Q_{50}, Q_{75}	Wealth Quartiles	LWS	−0.004, 0.027, 0.179	−0.005, 0.026, 0.182

TABLE 4
PARAMETERS CALIBRATED ENDOGENOUSLY

Parameter	Value	Description
Preferences		
φ	3.6	Bequest utility
$\beta_1, \beta_2, \beta_3$	0.952, 0.997, 0.952	Discount factors
χ	16.93	Disutility of work
Technology		
b	0.09	Borrowing limit
σ_ϵ	0.439	Variance of risk

in Brinca et al. (2016b). Due to the lack of panel data on individual incomes for European economies, which we could use to estimate the persistence of the idiosyncratic shock, ρ, we set it equal to the value used in Brinca et al. (2016b), who use U.S. data from the Panel Study of Income Dynamics (PSID). Setting $\rho = 0.335$ may at first appear low, but one should note that this estimation also includes a completely persistent fixed effect (ability).[29] The variance of the idiosyncratic income risk, σ_ϵ, is then calibrated to make the model match the variance of log wages in the data.

5.2. *Preferences and the Borrowing Limit.* The value of the Frisch elasticity of labor supply, η, has been much debated in the literature. We set it to 1, which is similar to that used in a number of recent studies; see, for example, Trabandt and Uhlig (2011) and Guner et al. (2016). The parameters, χ, governing the disutility of working an additional hour, φ, governing the utility of leaving bequests, the discount factors β_1, β_2, β_3, and the borrowing limit, b, are calibrated so that the model output matches the data. The corresponding data moments are average yearly hours, taken from the OECD Economic Outlook, the ratio of capital to output, K/Y, taken from the Penn World Table 8.0, and three wealth moments taken from the Luxembourg Wealth Study (LWS), namely the shares of wealth held by those between the 1st and 25th percentile, between the 1st and 50th percentile and between the 1st and 75th percentile. Finally, in order to have a realistic age profile of wealth, we also match the mean wealth held by those aged 75–80 relative to mean wealth in the whole population, from LWS.[30]

5.3. *Taxes and Social Security.* As described in the Appendix, we apply the labor income tax function in Equation (8), proposed by Benabou (2002). We use U.S. labor income tax data provided by the OECD to estimate the parameters θ_0 and θ_1 for different family types. To

[29] Brinca et al. (2016b) first estimate the age profile of wages. They then obtain the residuals, u_{it}, and estimate: $u_{it} = constant + a_i + \rho u_{it-1} + \epsilon_{it}$ where a_i is a fixed effect (ability). The Appendix contains further details and discussion of this estimation.

[30] Due to the small number of observations per cohort for most European countries, we match mean wealth held by those aged 75–80 in the U.S. economy.

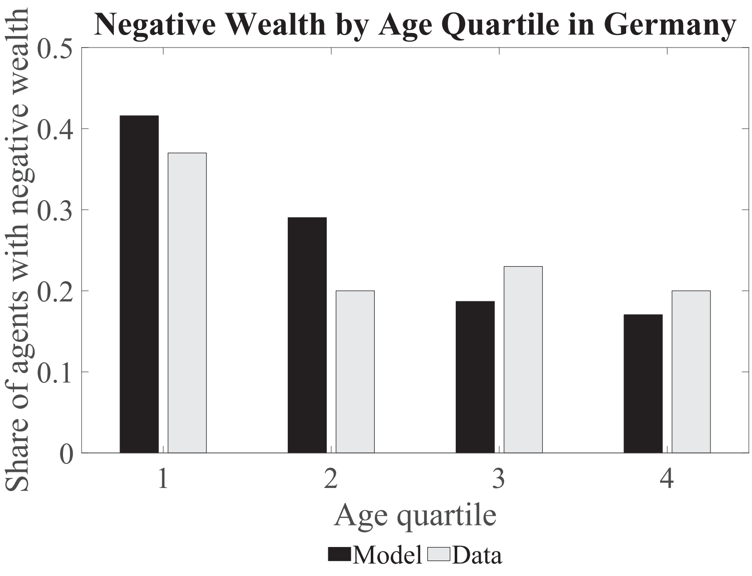


FIGURE 1

PERCENTAGE OF WORKING-AGE AGENTS WITH NEGATIVE WEALTH BY AGE QUARTILE IN THE MODEL (BLUE BARS) VERSUS EMPIRICAL OBSERVATIONS (YELLOW BARS), IN THE BENCHMARK ECONOMY GERMANY

obtain a tax function for the single individual households in our model, we take a weighted average of θ_0 and θ_1, where the weights are each family type's share of the population.[31]

For Germany, we estimate θ_0 and θ_1 to be 0.881 and 0.221, respectively. The employer social security contribution rate is set to 0.206 and the employee social security contribution rate to 0.21, taking the average tax rates between 2001 and 2007 from the OECD. Finally, consumption and capital tax rates are set to 0.233 and 0.155, respectively, following Trabandt and Uhlig (2011). The tax parameters for other countries is found in Table A21.

5.4. *Endogenously Calibrated Parameters.* To calibrate the parameters that do not have any direct empirical counterparts, φ, β_1, β_2, β_3, b, χ, and σ_ϵ, we use the simulated method of moments. We minimize the following loss function:

$$L(\varphi, \beta_1, \beta_2, \beta_3, b, \chi, \sigma_\epsilon) = ||M_m - M_d||, \tag{14}$$

where M_m and M_d are the moments in the data and in the model, respectively.

Given that we have seven parameters, we need seven data moments to have an exactly identified system. The seven moments we target in the data are the ratio of the average net asset position of households in the age cohort 75–80 relative to the average asset holdings in the economy, three wealth quartiles, the variance of log wages, and the capital to output ratio. All the targeted moments are calibrated with less than 2% of error margin, as displayed in Table 3. Table 4 presents the calibrated parameters. To illustrate that the model can also match some moments, not targeted in the calibration, Figure 1 compares the distribution of working-age agents with negative wealth by age decile in the model and in the data for the German benchmark economy. Since the fraction of borrowing constrained agents in the economy is important for our mechanism, it is reassuring that the model does quite well at matching the fraction of (working-age) agents with negative wealth by age.

[31] As we do not have detailed data for the population share of each family for European countries, we use U.S. family shares, as in Holter et al. (2019).

TABLE 5
MODEL SIMULATED LABOR SUPPLY ELASTICITY BY WEALTH

Wealth Quintile	Simulated Frisch Elasticity
1	0.36
2	0.43
3	0.48
4	0.49
5	0.51

6. INCOME INEQUALITY AND FISCAL CONSOLIDATION

In Section 3, we documented a strong empirical relationship between income inequality and the recessive impact of fiscal consolidation programs. This finding motivates the study of the impact of income inequality on fiscal consolidations in a structural model, which we explore in this section.

In the model, there are three sources of wage inequality: income risk, the permanent ability level, and the age profile of wages. We abstract from population growth and demographic differences across countries with respect to the relative sizes of each cohort.[32] There is an ongoing debate regarding whether income inequality is mainly due to differences between agents determined before their entry into the labor market or differences in the realization of income shocks during the life course. For example in the United States, Huggett et al. (2011) find that about 60% of the variance in lifetime earnings is due to initial conditions. This suggests that both initial conditions and market luck play an important role in generating the observed heterogeneity in the data.

In our structural model, we find that there is a link between income inequality, due to income risk, and the recessive impacts of fiscal consolidations. For inequality due to differences in initial conditions (ability and the age profile of wages in the model), this relationship is weak or nonexistent.

6.1. *Explaining the Mechanism: Income Risk, Precautionary Saving, and the Labor Supply Elasticity.* To understand the relationship between inequality and fiscal multipliers in our model, it may first be helpful to note that the intertemporal elasticity of substitution of labor (and consumption) is increasing in wealth.[33] Thus low-wealth agents will have a smaller labor supply response to future income changes (the case of austerity-based consolidation) as well as current price changes (the case of tax-based consolidation) but would, on the other hand, have a stronger labor response to changes in current transfers.[34] To verify that this is indeed the case in our model for current wage changes, we follow Blundell et al. (2016) who simulate the Frisch elasticity in a life cycle model by changing the wage rate by 1% at one age at a time (holding prices constant). They then average the response across age-groups. In our model, the response to a 1%, one-period wage change varies with wealth. In Table 5, we report the simulated "Frisch elasticity" in our model by wealth quintile.

As can be seen from the table, the model's labor supply elasticity is increasing in wealth. This experiment, where the wage is increased in the current period, is more similar to the model experiment with tax-financed consolidation (although there is also a future income effects in that experiment). The austerity-financed consolidation does not involve a current change in the wage rate, only a future income effect.

[32] For studies of the relevance of age structure for either fiscal policy or the effects of a credit crisis, see Basso and Rachedi (2017) and Antunes and Ercolani (2017).

[33] For empirical documentation of this fact, see Domeij and Floden (2006) for the relationship between wealth and IES of labor and Vissing-Jørgensen (2002), among others, for the relationship between wealth and the IES of consumption.

[34] See Brinca et al. (2016b) for an illustration of this case.

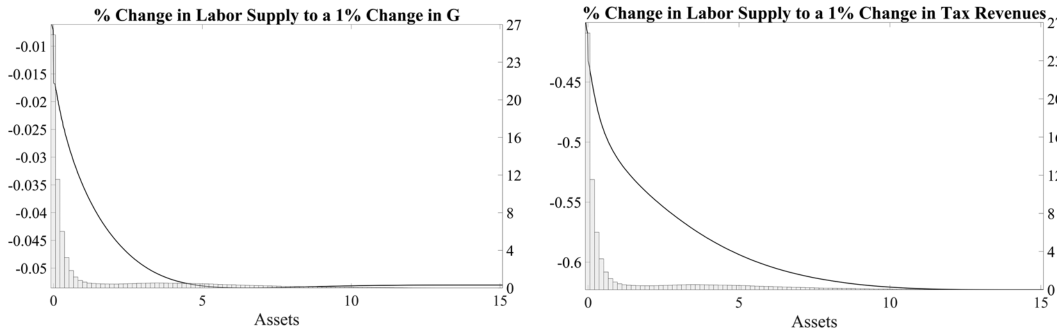


FIGURE 2

THE BLACK LINE STANDS FOR THE LABOR SUPPLY RESPONSE TO A 1% CHANGE IN G (LEFT PANEL) AND TO 1% INCREASE IN LABOR TAX REVENUES (RIGHT PANEL) BY ASSET LEVEL IN THE GERMAN BENCHMARK ECONOMY. THE RED BARS PRESENT THE MASS OF AGENTS IN PERCENT PER ASSET LEVEL

Now to understand how the full mechanism works in the context of our model, first consider the fiscal consolidation experiment where debt is reduced through a reduction in government spending. The decrease in government debt will gradually shift households' savings to physical capital, driving the capital-to-labor ratio up. The marginal product of labor in future time periods increases and this generates a positive shock to expected lifetime income for households, who in turn decrease their labor supply in the short run.[35] This effect also leads to a drop in output in the short run. However, given that productive capital increases during the transition to a new steady state, the economy will converge to a higher level of output in the long run. It should be noted that in our model, Ricardian equivalence does not hold for two reasons: first, given the overlapping generation structure of our model, older agents do not consider fiscal adjustments past their lifetime; second, given the incomplete markets structure of the model, agents cannot transfer resources across time and states of the world in response to the shocks, both from the lack of state-contingent contracts and the borrowing limit.

To understand the link between inequality and the initial drop in labor supply and output, note that the elasticity of labor supply to a shock to future income is smaller for credit-constrained and low-wealth agents (see the left panel of Figure 2). Constrained agents do not consider changes to their lifetime budget, only changes to their budget today (they have a low IES). Agents with low wealth levels are also less responsive to future income changes because they will be constrained in several future states of the world. An economy with high income inequality, arising from idiosyncratic productivity risk, has a smaller percentage of constrained and low-wealth agents, due to precautionary savings behavior, and a higher aggregate elasticity of labor supply with respect to our fiscal experiment, which causes a positive shock to future income. Therefore, a fiscal consolidation will be more recessive on impact in economies with high income inequality due to risk. In contrast, the variance of initial ability or the steepness of the age profile of wages will not affect the precautionary saving behavior of the agents, and changing the variance of ability or the slope of the age profile will have little to no impact on the number of credit-constrained agents.

In the case of consolidation through increased labor income taxes, the mechanism through which inequality matters share some of the features described above, namely the negative

[35] Note that in the austerity-case, the mechanism works almost solely through changes in prices. If the change in debt does not have an impact on prices, as can arguably be the case of in a small open economy, then the impact of austerity-based consolidation is greatly diminished, reflecting only the income effect from a smaller debt service in the future. However, there are several reasons to believe that even in small open economies debt consolidation would affect prices. For example, Chakraborty et al. (2017) document that, across countries, a disproportionate amount of commercial debt is held by nationals. And in the presence of some form of home-equity bias, a reduction in government debt could have real effects in the economy. As another example, foreigners hold only 22% of total sovereign debt from the Italian Government, and 7% in the case of Japan—see Gros (2019).

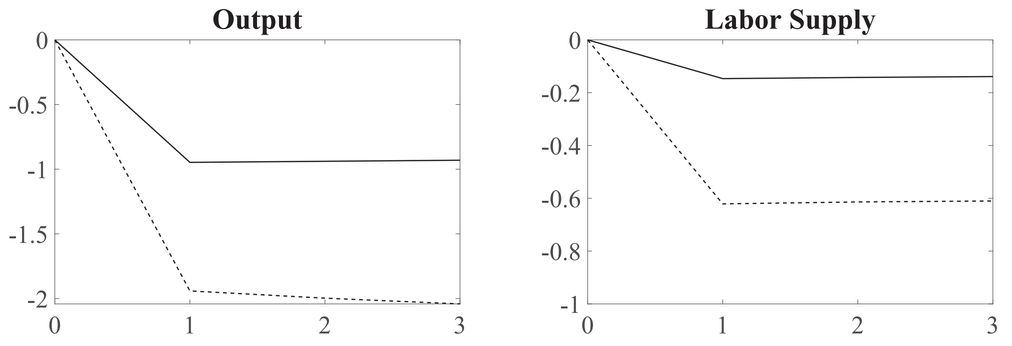


FIGURE 3

LABOR TAX CONSOLIDATION: OUTPUT CUMULATIVE MULTIPLIER (LEFT PANEL) AND LABOR SUPPLY CUMULATIVE MULTIPLIER (RIGHT PANEL) IN THE FIRST THREE PERIODS IN GERMANY (DASHED LINE) AND CZECH REPUBLIC (SOLID LINE)

income effect on labor supply today, through higher future wages (the total income effect from consolidation could be positive or negative depending on whether the wage increase dominates the increase in future taxes). However, the tax also induces a negative substitution effect on wages today, both for constrained and unconstrained agents. Due to the negative substitution effect, unconstrained agents would like to reduce labor supply and use their savings or borrow to avoid a large drop in consumption. However, for constrained agents, who do not consider their life-time budget but only their budget today, this is not an option. To avoid a big drop in consumption today, they must work more today and thus their labor supply response to a fall in the net wage rate is smaller.

6.2. *Illustrating the Mechanism: Comparing Fiscal Consolidation in Germany and the Czech Republic.* To illustrate the impact of differences in income inequality, we first compare the effects of consolidation in Germany and in the Czech Republic—two European countries on the opposite side of the spectrum in terms of wage inequality. Germany has the second highest variance of log wages, 0.354, and Czech Republic has the lowest value at 0.174. Although these two countries differ along several dimensions, the primary motivation for comparing Germany and the Czech Republic is due to their differences in wage inequality, idiosyncratic risk, and the percentage of constrained agents. In the Czech Republic, the calibrated variance of the idiosyncratic risk is 0.145 and the percentage of constrained agents is 7.39%, whereas Germany has a higher variance of risk, 0.439, and a lower percentage of constrained agents, 3.41%. We find what our mechanism suggests that the output multiplier following the unanticipated fiscal consolidation shock is larger in Germany than in Czech Republic.

In Figures 3 and 4, we plot the cumulative output multiplier and labor supply response to labor tax and government spending consolidations, respectively, for the two countries. Both the labor supply responses and the output multipliers are significantly larger in the German economy, where wage inequality is higher. As Germany has a smaller share of constrained and low-wealth agents, the output drop is more pronounced. One should also note that the consolidation through increased labor income taxes causes deeper recessions than the consolidation financed by a reduction in government spending. This is consistent with the results in Alesina et al. (2017).

6.3. *Inequality: Variance of Risk versus Variance of Ability versus Age Profiles.* Next, we perform three experiments in our German benchmark economy to verify the mechanism described above. We focus on understanding the role of the different parameters that drive wage inequality in our model, σ_ϵ, σ_a, and γ_1, γ_2, γ_3. These parameters govern the variance of idiosyncratic wage shocks, the variance of permanent ability, and the shape of the age profile of

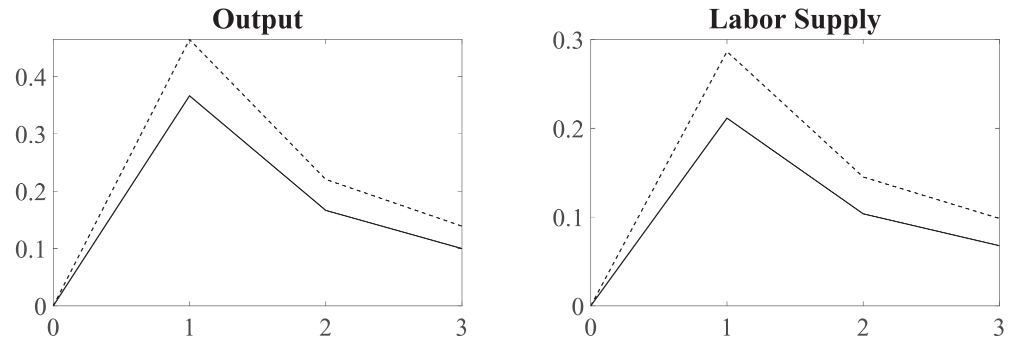


FIGURE 4

GOVERNMENT SPENDING CONSOLIDATION: OUTPUT CUMULATIVE MULTIPLIER (LEFT PANEL) AND LABOR SUPPLY CUMULATIVE MULTIPLIER (RIGHT PANEL) IN THE FIRST THREE PERIODS IN GERMANY (DASHED LINE) AND CZECH REPUBLIC (SOLID LINE)

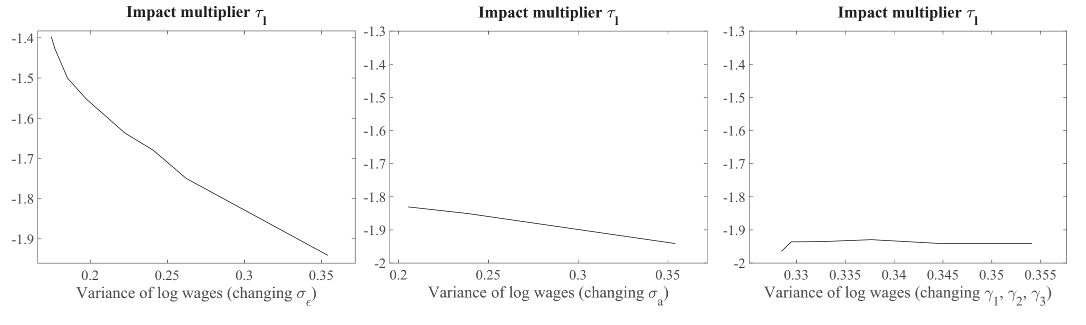


FIGURE 5

IMPACT MULTIPLIER FOR THE LABOR TAX CONSOLIDATION IN THE BENCHMARK MODEL FOR GERMANY WHEN CHANGING THE VARIANCE OF RISK (LEFT PANEL), THE VARIANCE OF ABILITY (MIDDLE PANEL), AND THE AGE PROFILE OF WAGES (RIGHT PANEL)

wages. We find that the correlation between wage inequality and fiscal multipliers that we documented in the empirical section can only be explained by differences in idiosyncratic risk and not by predetermined differences in ability or in the age profile of wages. We perform three different experiments:

(1) We gradually change $\mathrm{Var}(\ln w)$ in the benchmark model calibrated to Germany through altering the variance of innovations to the stochastic income process, σ_ϵ^2, moving from zero to the highest value in our sample of countries.
(2) We gradually change $\mathrm{Var}(\ln w)$ in the benchmark model calibrated to Germany, by altering the variance of ability, σ_a^2, rising from zero to the value used throughout in our calibration exercises (see Table A23).
(3) We gradually change $\mathrm{Var}(\ln w)$ in the benchmark model calibrated to Germany, by multiplying the age profile of wages, governed by γ_1, γ_2, γ_3, by a *Scalar*, ranging from 0 to 1.

In all cases, we adjust γ_0 by a constant to guarantee that average productivity in the economy stays unchanged. Then for each value of σ_u, σ_a, and the *Scalar*, we perform our two fiscal consolidation experiments: (i) consolidation through reductions in government spending and (ii) consolidation through increases in the labor income tax.

In Figure 5, we plot the impact multiplier in the experiment with fiscal consolidation through labor income taxes for different values of σ_ϵ, σ_a, and the *Scalar*. In the left panel, we observe that the fiscal multiplier is very sensitive to changes in income risk. When we change

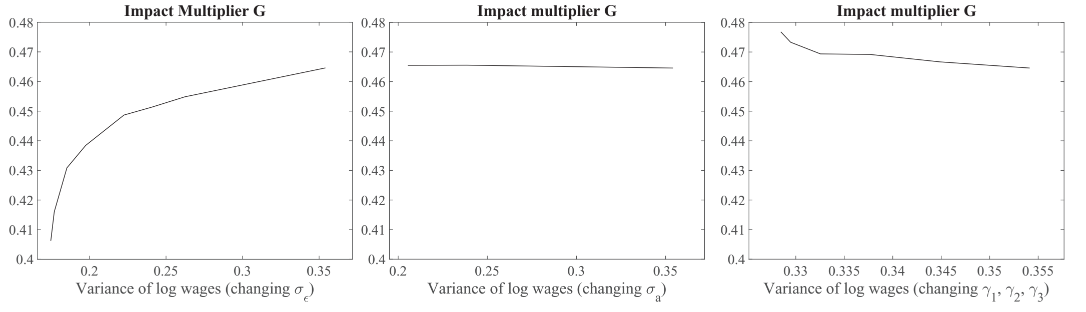


FIGURE 6

IMPACT MULTIPLIER FOR THE CONSOLIDATION THROUGH GOVERNMENT SPENDING IN THE GERMAN BENCHMARK ECONOMY WHEN CHANGING THE VARIANCE OF RISK (LEFT PANEL), THE VARIANCE OF ABILITY (MIDDLE PANEL), AND THE AGE PROFILE OF WAGES (RIGHT PANEL)

the variance of the innovations to the idiosyncratic shock, ϵ, from 0 to 0.45 the impact multiplier falls from about -1.40 to -1.95. In the middle and right panels, we observe that it is relatively inelastic with respect to changes in ability and the steepness of the age profile of wages.[36]

The experiment with consolidation through government spending generates similar results. In the left panel of Figure 6, we observe that as we change the variance of the innovations to the idiosyncratic shock, governed by σ_ϵ, from 0 to 0.45 the impact multiplier increases from about 0.41 to 0.47. In the middle and the right panels of the figure, we observe that the changes in the multiplier induced by changing the variance of ability and the steepness of the age profile of wages are small. We conclude that only through changes in income risk can we generate a positive relationship between the impact of fiscal consolidation programs and income inequality.

The analysis in Figures 5 and 6 covers changes in risk that go from zero to the highest value obtained in our calibration of the model to 13 different European countries. In our calibration exercise, the lowest value of the variance of risk was obtained for Greece and equal to 0.12 and the highest was equal to 0.5, for France. One should note that the relative magnitude of changes in the multiplier induced by changing the risk is larger for tax-based than for spending-based consolidation. Going from the lowest to the highest level of risk implies a 30% increase in the impact multiplier for the tax-based consolidation and an 8% increase in the impact multiplier for the spending-based consolidation. As mentioned before, it is worth noting that the actual consolidations studied in Section 3 include changes to both taxes and spending.

In Figure 7, we verify our hypothesis about the relationship between income risk and the fiscal consolidation multipliers stemming from the fact that economies with higher income risk have a lower share of credit-constrained agents. In the left panel of the figure, we document a strong, negative relationship between the variance of risk and the proportion of credit-constrained agents in the economy. In the middle panel, we see that changing the variance of ability does not affect the share of agents with liquidity constraints, as we anticipated. A steeper age profile of wages leads to more liquidity-constrained agents (as one would expect[37]) but the effect is very weak compared to the impact of income risk.

In Figure 8, we illustrate the relationship between the share of agents with liquidity constraints and the impact multiplier, for both spending-based and tax-based fiscal consolidation,

[36] Germany has one of the steepest age profiles in our sample of countries. We therefore let the scalar go from 0 to 1, capture the effect of going from a steep age profile to a completely flat age profile.

[37] With a steeper age, profile agents will save less early in the life cycle.

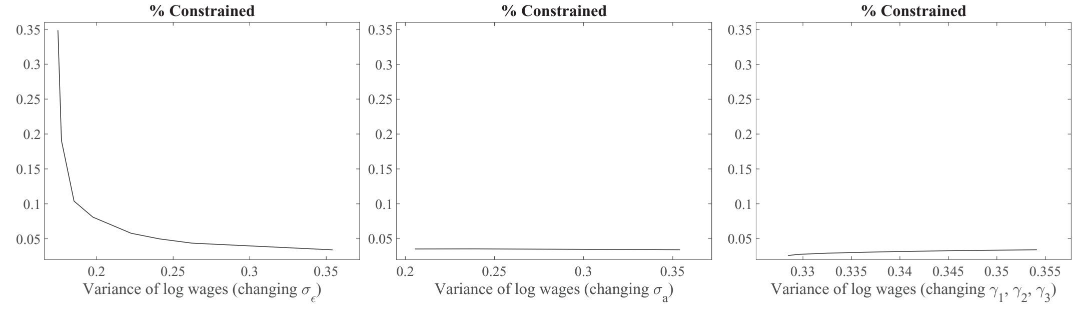


FIGURE 7

SHARE OF CREDIT-CONSTRAINED AGENTS IN THE GERMAN BENCHMARK ECONOMY WHEN CHANGING THE VARIANCE OF RISK (LEFT PANEL), THE VARIANCE OF ABILITY (MIDDLE PANEL), AND THE AGE PROFILE OF WAGES (RIGHT PANEL)

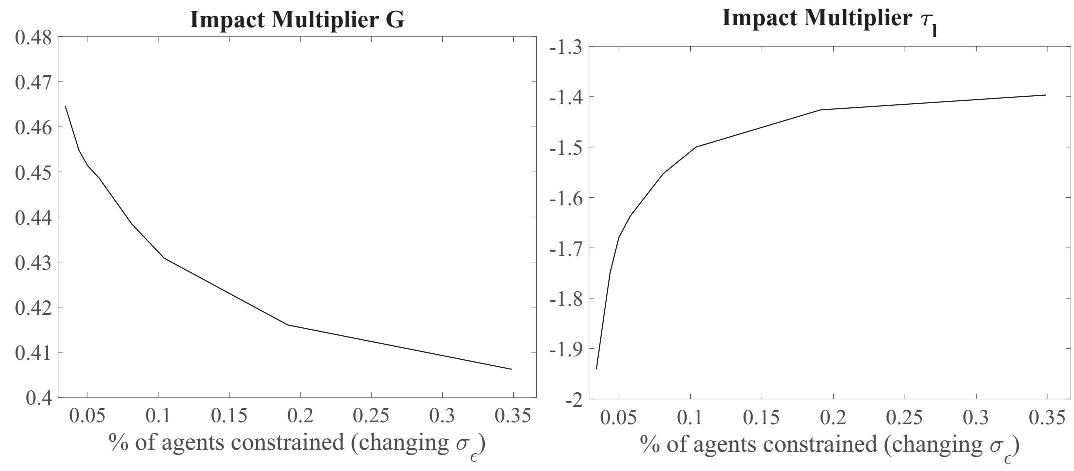


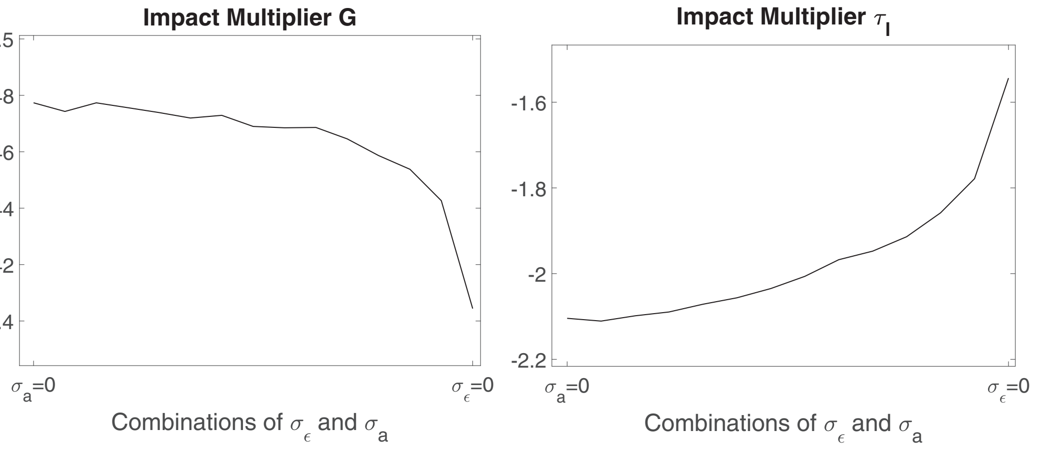
FIGURE 8

IMPACT MULTIPLIER FOR THE G-CONSOLIDATION (LEFT PANEL) AND FOR THE τ_l-CONSOLIDATION (RIGHT PANEL) PLOTTED AGAINST THE SHARE OF CREDIT-CONSTRAINED AGENTS IN THE GERMAN BENCHMARK ECONOMY, WHEN DECREASING THE VARIANCE OF RISK

as we change income risk. We observe that there is a strong negative relationship between the share of credit-constrained agents and the fiscal consolidation multipliers.

Finally, as a last robustness test to verify that the relationship between inequality and fiscal multipliers comes from the variance of risk, we conduct the following experiment: we keep wage inequality constant by choosing different combinations of risk and ability, going from one extreme, where all wage inequality (except from the age profile) is due to the variance of risk, to the other extreme, where wage inequality is fully explained by variance of ability. Figure A4 shows that the multiplier is largest when all inequality is explained by income risk and smallest when all inequality is explained by the variance of ability, for both tax-based and expenditure-based consolidations.

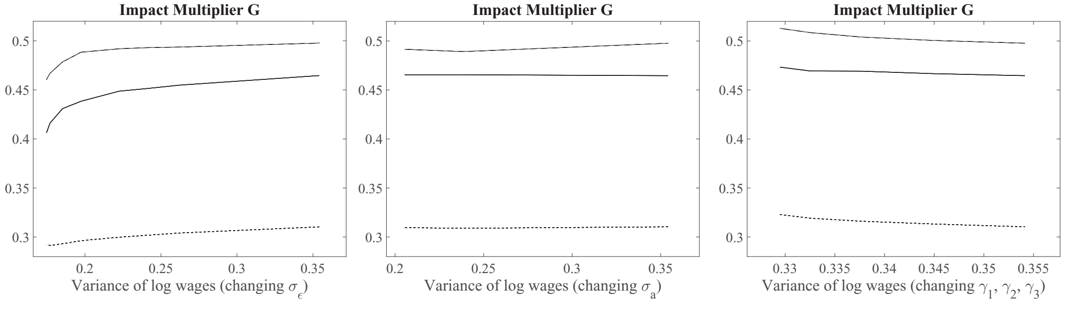
6.4. *Robustness: Risk Aversion.* As our mechanism hinges on wealth effects and future income changes, we produce robustness tests with respect to the value of the risk aversion parameter σ and the relative importance of income risk, the variance of ability, and the age profile of wages for the fiscal multipliers. Figures A5 and A6 present the results for these robustness exercises. We replicate Figures 5 and 6 with the following alternate assumptions on the risk aversion parameter, $\sigma = 0.5$ and $\sigma = 4$. We recalibrate the remaining parameters to



NOTES: The variance of ability is increasing from left to right. When the variance of ability is zero, all inequality is due to risk, which is calibrated to keep the variance of log wages at the benchmark value. When the variance of risk is zero, all inequality is due to the variance of ability, which is calibrated to match the variance of log wages.

FIGURE A4

IMPACT MULTIPLIER FOR THE G-CONSOLIDATION (LEFT PANEL) AND FOR THE τ_l-CONSOLIDATION (RIGHT PANEL) FOR DIFFERENT COMBINATIONS OF VARIANCE OF ABILITY AND RISK ON THE X-AXIS, KEEPING VARIANCE OF LOG WAGES CONSTANT



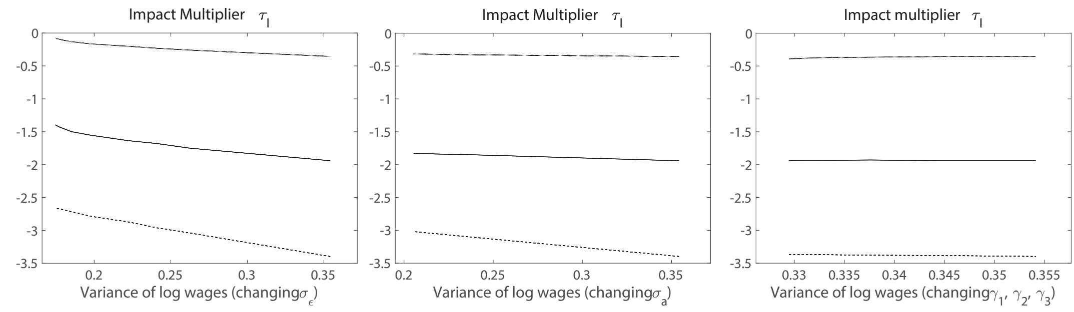
NOTES: The dashed line illustrates the results when $\sigma = 0.5$, the solid line for the benchmark value of $\sigma = 1.2$, and the dotted line for $\sigma = 4$.

FIGURE A5

IMPACT MULTIPLIER FOR THE CONSOLIDATION THROUGH GOVERNMENT SPENDING IN THE GERMAN BENCHMARK ECONOMY WHEN CHANGING THE VARIANCE OF RISK (LEFT PANEL), THE VARIANCE OF ABILITY (MIDDLE PANEL), AND THE AGE PROFILE OF WAGES (RIGHT PANEL)

guarantee that we still match the same data moments from the reference calibration. Figure A5 presents the results for the spending consolidation when changing the variance of risk (left panel), the variance of ability (middle panel), and the age profile of wages (right panel), for the benchmark economy with $\sigma = 1.2$ (solid line), for $\sigma = 0.5$ (dashed line) and for $\sigma = 4$ (dotted line). The results are qualitatively, and to a large extent quantitatively, robust to any of the three values of σ, with the variance of risk being the income inequality source that generates a positive correlation between income inequality and fiscal multiplier. The value of σ has a significant impact on the level of the multiplier but it is still increasing in income risk. Figure A6 presents the same exercise for the labor tax consolidation. Again, risk is the main driver of the positive correlation between the multiplier and income inequality, independently of the value of σ.

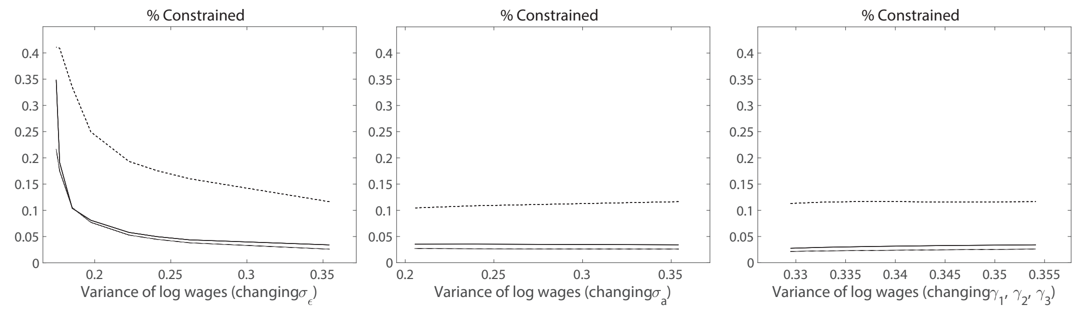
Figure A7 illustrates the mechanism behind the positive correlation between the multiplier and income inequality. Similar to the benchmark economy, with $\sigma = 1.2$, the percentage of constrained agents is invariant to the variance of ability and to the age profile of wages for any



NOTES: The dashed line illustrates the results when $\sigma = 0.5$, the solid line for the benchmark value of $\sigma = 1.2$, and the dotted line for $\sigma = 4$.

FIGURE A6

IMPACT MULTIPLIER FOR THE CONSOLIDATION THROUGH GOVERNMENT SPENDING IN THE GERMAN BENCHMARK ECONOMY WHEN CHANGING THE VARIANCE OF RISK (LEFT PANEL), THE VARIANCE OF ABILITY (MIDDLE PANEL), AND THE AGE PROFILE OF WAGES (RIGHT PANEL)



NOTES: The dashed line illustrates the results when $\sigma = 0.5$, the solid line for the benchmark value of $\sigma = 1.2$, and the dotted line for $\sigma = 4$.

FIGURE A7

PERCENTAGE OF CONSTRAINED AGENTS IN THE BENCHMARK MODEL FOR GERMANY WHEN CHANGING THE VARIANCE OF RISK (LEFT PANEL), THE VARIANCE OF ABILITY (MIDDLE PANEL), AND THE AGE PROFILE OF WAGES (RIGHT PANEL)

values of σ. For any value of σ, the constrained and low-wealth agents still have the lowest labor supply response to the shock and the mechanism still hinges on the steady state percentage of constrained agents. With the mass of constrained agents decreasing with the variance of risk, but invariant to the variance of ability and the age profile of wages, it changes in the variance of risk that generate variations in the fiscal multiplier.

7. CROSS-COUNTRY ANALYSIS

In the previous section, we demonstrated that our model is able to reproduce the empirical relationship between income inequality and fiscal multipliers, through variation in income risk. In this section, we perform a cross-country analysis to show that this mechanism is strong enough to matter quantitatively. We calibrate our model to match a wide range of different country characteristics, where, in addition to the distributions of income and wealth, we match data on taxes, social security, and government debt. We show that even when introducing substantial country heterogeneity, we are able to reproduce the cross-country relationship between both tax- and spending-based fiscal consolidation and income inequality. In Subsections 7.1 and 7.2, we verify that our results are robust to recalibrating all countries using the distribution of liquid wealth instead of net wealth, and to assuming that all countries are small open economies.

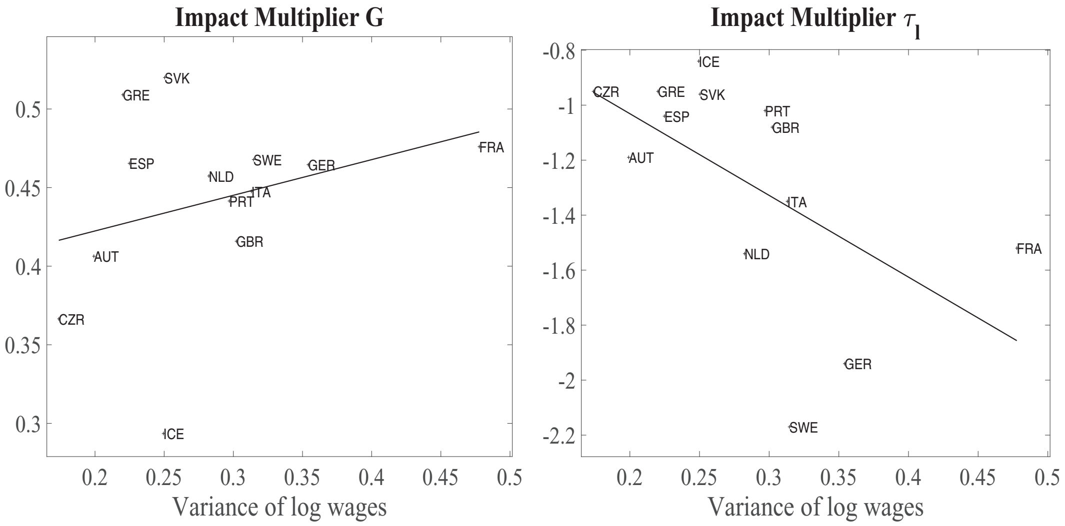


FIGURE 9

IMPACT MULTIPLIER AND VAR(LN(W)). THE LEFT PANEL DISPLAYS THE CROSS-COUNTRY DATA FOR A CONSOLIDATION IMPLEMENTED BY DECREASING G (CORRELATION COEFFICIENT 0.35, $p = 0.25$), AND THE RIGHT PANEL DISPLAYS THE CROSS-COUNTRY DATA FOR A CONSOLIDATION IMPLEMENTED BY INCREASING THE LABOR TAX (CORRELATION COEFFICIENT −0.60, $p = 0.03$)

The model is calibrated to 13 European countries[38] using country-specific age profiles of wages, keeping the variance of the permanent ability fixed and changing the variance of the idiosyncratic shock to match the variance of log wages in the data.[39] Tables A2 and A21 summarize the wealth distribution, the other country-specific data used to calibrate the model, and the country-specific parameters estimated outside of the model. Table A22 summarizes the country-specific parameters estimated through the simulated method of moments, as described in Section 5. Parameters kept constant for all the countries are summarized in Table A23.

Figure 9 reveals that our model is able to reproduce the cross-country empirical relationship between income inequality and the impacts of fiscal consolidation: countries with higher income inequality experience larger output drops on impact, both for tax- and spending-based consolidations. These effects are large and economically meaningful, in particular for tax-based consolidations. Using the coefficient found when regressing the multiplier on income inequality, we find that the response between the country with the lowest income inequality (Czech Republic) and the highest (France) leads to a 90% increase in the tax-based multiplier. One should also note that in our model exercises, tax-based consolidations produce deeper recessions across countries than spending-based consolidations, in line with most literature.

In the previous section, we argued that the mechanism through which higher income risk translates into larger multipliers is through changes in the share of credit-constrained agents. In Figure 10, this relationship is documented for the 13 economies for which we calibrate the

[38] For this exercise, we use only countries that actually underwent fiscal consolidation processes after 2009. Compared to Blanchard and Leigh (2013), we are forced to exclude Belgium, Cyprus, Denmark, Ireland, Malta, Norway, Poland, Romania, and Slovenia due to data limitations. The results in Subsection 3.1 are, however, robust to considering only these 13 countries. See Table A13.

[39] A recent working paper by Halvorsen et al. (2019) using register data from Norway and United States establishes that there are substantial differences in income risk between these two countries.

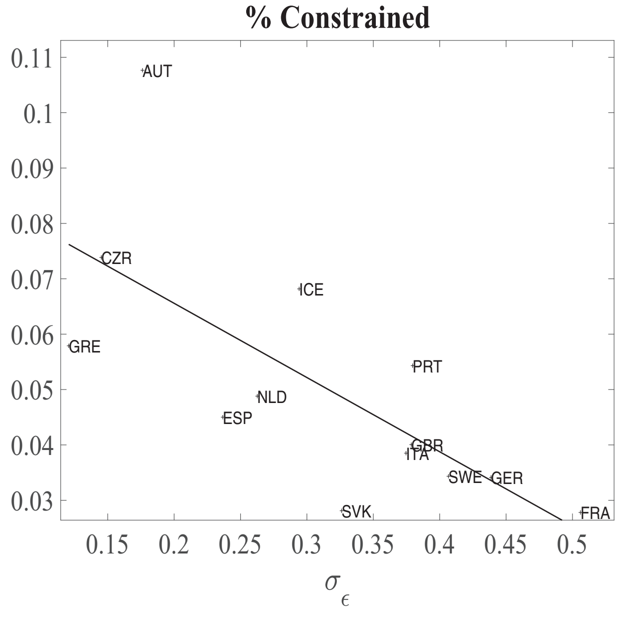


FIGURE 10

PERCENTAGE OF AGENTS CONSTRAINED IN THE Y-AXIS AND VARIANCE OF IDIOSYNCRATIC RISK ON THE X-AXIS. CORRELATION COEFFICIENT OF -0.73 AND P-VALUE OF 0.0

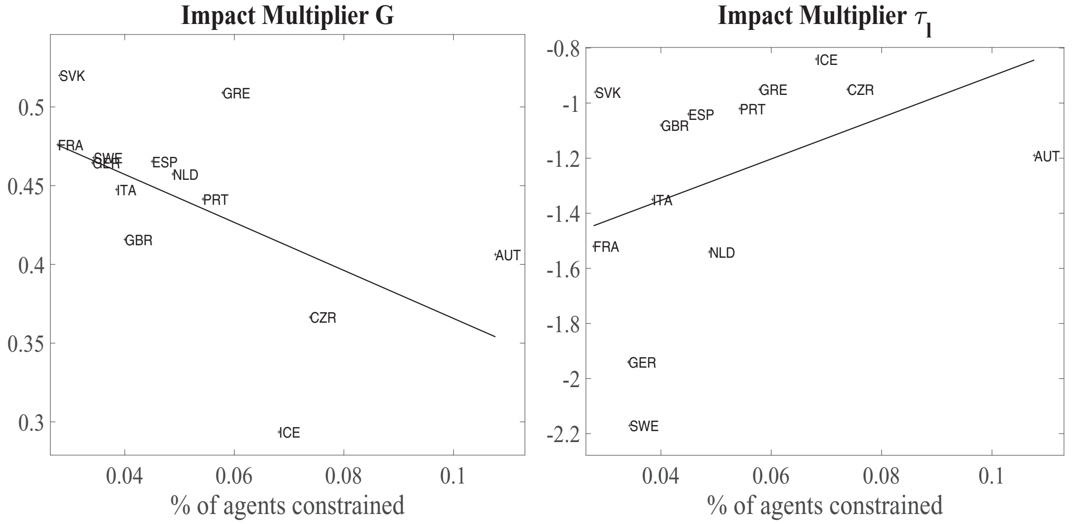
model. Countries with a higher standard deviation of the innovations to idiosyncratic income risk, σ_ϵ, have a smaller share of constrained agents.[40]

As argued before, the labor supply of constrained agents is less elastic with respect to the fiscal shock, and the larger the percentage of constrained agents, the smaller the multiplier. In Figure 11, this relationship is documented across countries. Countries with a larger share of liquidity-constrained agents experience a smaller output drop for both spending- and tax-based consolidations.

7.1. *Robustness: Recalibration with Liquid Wealth Moments.* One may wonder if income risk continues to matter for a different distribution of wealth. A recent literature has emphasized the importance of liquid wealth for the marginal propensity to consume (and work).[41]

[40] Note that for our mechanism, it is ultimately *post*-tax income inequality that matters and the heterogeneity in the progressivity of each country tax system, the level of transfers, and social security add substantial noise to the regression above, which describes pre-tax wage income inequality. We run the regression on pre-tax wages to be consistent with the evidence we provide in the empirical section and still find a strong and significant relationship. Still, the results are robust to using post-tax labor income inequality as can be seen in Figure A8. We also compute the correlation between the fiscal multipliers and the progressivity of the tax code across countries in Figure A9. It has the expected negative sign in the case of the G-consolidation but one should have expected a positive correlation in the case of tax-based consolidation, at least thinking about the effect of tax progressivity on income inequality. Tax progressivity is, however, strongly correlated with other potentially important factors that we let vary across countries such as tax level and social security. A full analysis of the interaction between welfare state policies and fiscal consolidation is beyond the scope of this article.

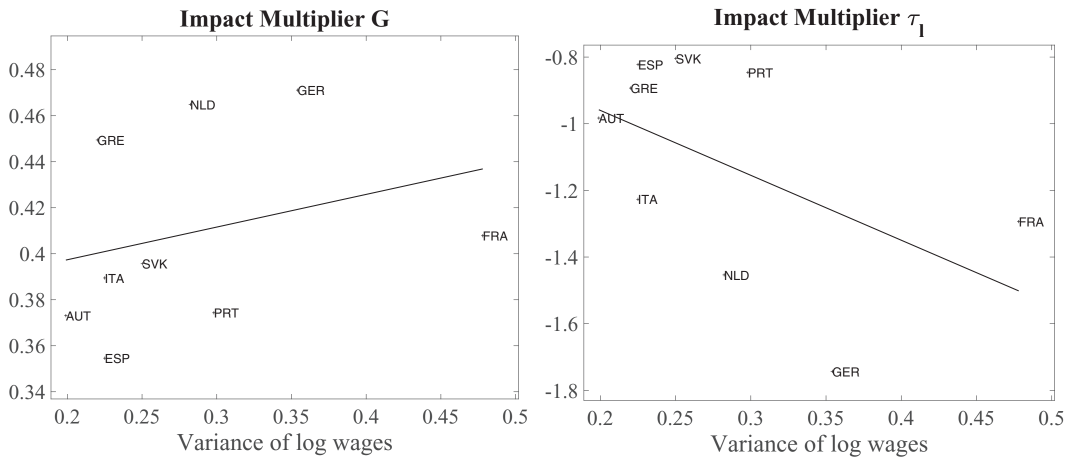
[41] Carroll et al. (2014) measure marginal propensities to consume for a large panel of European countries, and then calibrate a model for each country using net wealth and liquid wealth. The authors find that the higher the proportion of financially constrained agents in an economy, the higher the consumption multiplier. Kaplan and Violante (2014) propose a model with two types of assets that provide a rationale for relatively wealthy agents' choice of being credit constrained. In a context of portfolio optimization with one high-return illiquid asset and one low-return liquid asset, relatively wealthy individuals may end up credit constrained. Kaplan et al. (2014), using micro data from several countries, then argue that the percentage of financially constrained agents can be well above what is typically the outcome of models where very few agents have their wealth tied up in illiquid assets.



NOTES: On the left panel, we have the cross-country data for a consolidation done by decreasing G (correlation coefficient -0.68, $p = 0.01$), whereas on the right panel, we have the cross-country data for a consolidation done by increasing the labor tax (correlation coefficient 0.55, $p = 0.06$)

FIGURE 11

IMPACT MULTIPLIER AND PERCENTAGE OF AGENTS CONSTRAINED

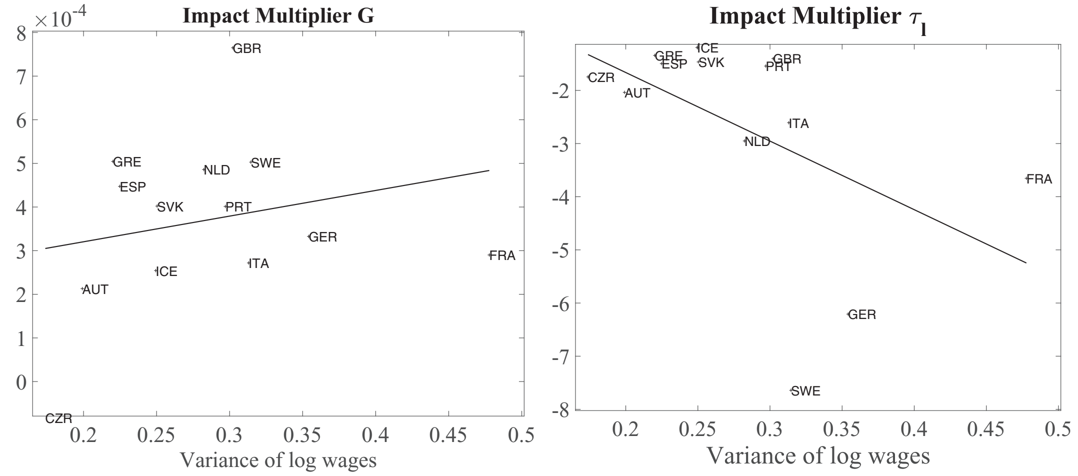


NOTES: On the left panel, we have the cross-country data for a consolidation done by decreasing G (correlation coefficient 0.4500; $p = 0.2298$), whereas on the right panel we have the cross-country data for a consolidation done by increasing the labor tax (correlation coefficient -0.3833; $p = 0.3125$).

FIGURE A10

IMPACT MULTIPLIER AND VAR(LN(W)) IN A MODEL CALIBRATED TO MATCH THE LIQUID WEALTH DISTRIBUTION

Our framework does not separate wealth into multiple assets, as in Kaplan et al. (2014), however, we can examine the distribution of *liquid* assets across countries from the Household Finance and Consumption Survey (HFCS)—also used by Carroll et al. (2014)—and recalibrate the model for our sample of countries, holding the other calibration targets constant. The distribution of liquid wealth is much more unequal than the distribution of net wealth and such a calibration delivers more credit-constrained agents. Figure A10 shows the correlation between income inequality and the fiscal multipliers in our two consolidation exercises. Income risk continues to influence the pattern of response to consolidations under the newly calibrated



NOTES: On the left panel, we have the fiscal multiplier across countries for the G consolidation, in a setting where wages and interest rate are constant. On the right panel, we have the fiscal multipliers across countries for the labor tax experiment in the same constant wages and interest rate environment.

FIGURE A13

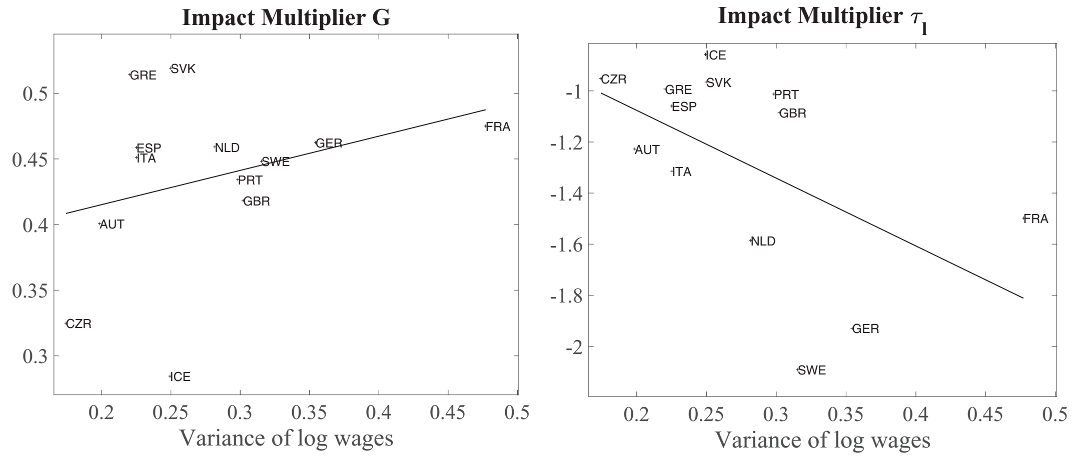
IMPACT MULTIPLIER AND VAR(LN(W)) IN A PARTIAL EQUILIBRIUM MODEL

wealth distribution (which we match through the heterogeneous discount factors in our calibration). This is because the income risk affects the number of credit-constrained agents, the degree of precautionary savings, and the distribution of wealth just above the borrowing constraint.[42]

7.2. *Robustness: Open Economy.* Many of the countries in our sample can be thought of as small open economies. Therefore, we test the extent to which our mechanism works if the wage and interest rate are set in the international market and invariant to the fiscal consolidation shocks. Under this alternative assumption, we derive results that are qualitatively consistent with the general equilibrium model for both taxation and government spending consolidation experiments, see Figure A13. Specifically, in partial equilibrium, there is a positive correlation between income inequality and the fiscal multipliers for both tax-based and austerity-based consolidation. However, the multipliers become (much) smaller with respect to austerity-based consolidation, and larger for labor-tax-based consolidation. In the spending-based consolidation, the average size falls from 0.44 to 0.0004 under small open economy assumptions. Although the opposite occurs in the tax-based consolidation, with the average multiplier going from -1.27 to -2.71. The mechanisms driving spending and tax-based consolidations are future income effects and a current substitution effect (for tax-based consolidation). With constant wages, the future income effect becomes considerably smaller (only stemming from the change in the lump-sum transfer once the consolidation is achieved, given the smaller debt service), whereas the current substitution effect is amplified (the current wage is not responding to the fall in labor). This explains the smaller multipliers for the expenditure-based consolidation and the larger ones for the labor tax experiment.[43] The

[42] Figure A11 displays the correlation between the share of top 20% over share of bottom 20% of the liquid wealth distribution and the consolidation multipliers, whereas Figure A12 displays the correlation between the share of top 20% over share of bottom 20% of net wealth distribution and the multipliers in our benchmark calibration. In the two different calibration strategies and for both experiments, we see the correct correlation sign between wealth inequality and multipliers (more wealth inequality implies a smaller response to shocks). Despite this, the correlation between income risk and the multipliers is evident in both cases.

[43] There are several reasons to believe that even in small open economies, the mechanism we describe could have an impact on prices. For example, we find that a disproportionate amount of government debt is held by nationals,



NOTES: The left panel displays the cross-country data for the G consolidation multiplier (correlation coefficient 0.31; $p = 0.30$), whereas the right panel displays cross-country data for the labor tax consolidation (correlation coefficient -0.60, $p = 0.03$).

FIGURE A14

IMPACT MULTIPLIER FOR THE EXPERIMENT WHERE RISK AND ABILITY EACH EXPLAIN 50% OF THE VARIANCE OF LOG WAGES AND VAR(LN(W))

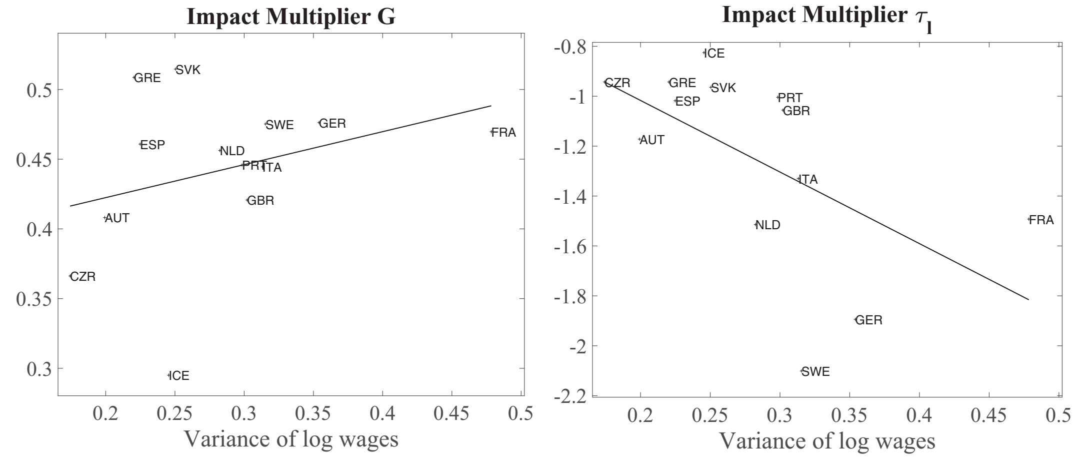
reason that the mechanism is still quantitatively strong, in the case of tax-based consolidation, is that unconstrained agents have some savings or borrowing capacity. They can use this capacity to smooth consumption when the unexpected tax shock hits, and they will reduce labor supply more than constrained agents, who have to work hard today to avoid a large drop in consumption.

7.3. *Robustness: Risk versus Ability as Determinants of Wages.* There is an ongoing debate on how much of observed wage inequality is due to labor market risk, and one may wonder if our results would be robust to having the cross-country differences in income inequality (after controlling for age profiles) not be fully explained by the risk component. In this subsection, we propose a robustness exercise where we let the variance of ability and of risk each explain 50% of the variance of log wages in each country, after controlling for the age profile of wages. We then recalibrate the remaining parameters to hit the same data targets as before. The results are presented in Figure A14 and illustrate that our mechanism is robust to this new calibration strategy, with the correlation between income inequality and multiplier being similar to the benchmark case, both qualitatively and quantitatively.

To develop an understanding for why the results are similar, Table A20 reports the σ_u, σ_a, and the percentage of constrained agents, both in the benchmark calibration and when we set both risk and ability each to explain 50% of the residual variance of log wages. We observe that the differences in the percentage of constrained agents are relatively small between the two calibrations, which explains why the results hold both qualitatively and quantitatively.

7.4. *Robustness: Persistence and Variance of Innovations to Wage Shocks.* Finally, we perform a robustness test with respect to the value of ρ, which is set to 0.34 in the bench-

and in the presence of some form of home-equity bias, a reduction in government debt could lead to the crowd-in effect on productive capital that would impact aggregate prices. Consistent with this is the fact that foreigners hold only 22% of total sovereign debt from the Italian Government, and 7% in the case of Japan—see Gros (2019). Another factor is timing. Perhaps in an open economy consolidation would affect short-run prices, before they adjust back to world market prices.



NOTES: On the left panel, we have the cross-country data for a consolidation done by decreasing G (correlation coefficient 0.35; $p = 0.24$), whereas on the right panel we have the cross-country data for a consolidation done by increasing the labor tax (correlation coefficient -0.71; $p = 0.01$).

FIGURE A15

IMPACT MULTIPLIER FOR THE EXPERIMENT WHEN $\rho_u = 0.8$ AND VAR(LN(W))

mark economy.[44] We instead set $\rho = 0.8$, a value similar to what other authors such as Chang and Kim (2006) have estimated, while recalibrating σ_ϵ to match the variance of log wages across the different countries in our sample. We also recalibrate the remaining parameters to match the remaining calibration targets. Figure A15 displays the cross-country correlation between the fiscal multipliers and the variance of log wages for both spending and labor tax consolidations for this new calibration. The results are similar to the ones presented in Section 7, with the correlation between the multiplier and income inequality being positive and significant.

It is perhaps a bit surprising that the results are so similar across the two calibration strategies. However, one should remember that the unconditional variance of the income risk process is similar across the two experiments. Recall that the unconditional variance of u, determined by the AR(1) process, is given by

$$\mathrm{Var}(u) = \frac{\sigma_\epsilon^2}{1 - \rho^2}.$$

By increasing ρ, and consequently decreasing σ_ϵ to match the variance of log wages, the unconditional variance of income risk remains the same across countries. At the same time, other model parameters are also being recalibrated to keep the same wealth distribution and this may explain why the difference is so small compared to the benchmark experiment.

8. EMPIRICAL VALIDATION OF THE MECHANISM

In Section 3, we established that income inequality amplifies the recessive effects of fiscal consolidations. In Section 6, we study a mechanism that leads to this amplification effect: labor supply responds stronger in countries with higher income inequality, leading to a larger output drop. In Section 8, we present two pieces of empirical evidence that support our mechanism. First, we use the fact that household debt and the share of credit-constrained agents are strongly correlated in our benchmark economy. If our mechanism is correct, the output

[44] See the Appendix for a detailed discussion of why our estimation procedure for ρ produces a lower value than what is typically found in the literature.

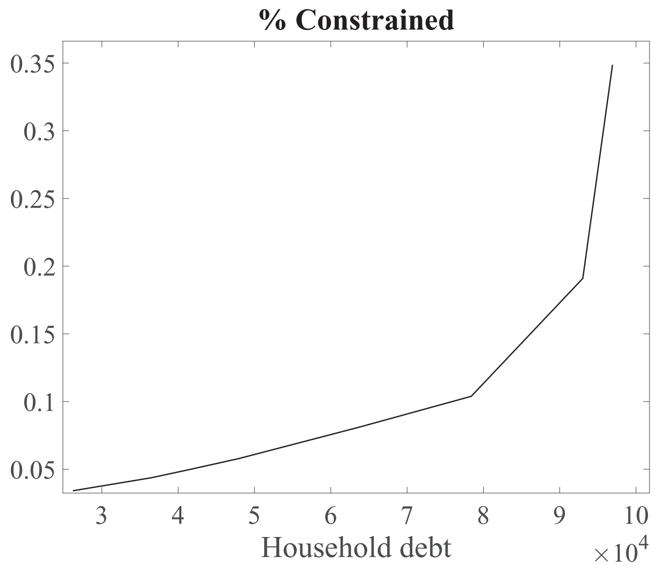


FIGURE 12

THE FRACTION OF LIQUIDITY-CONSTRAINED AGENTS (Y-AXIS) AND HOUSEHOLD DEBT LEVEL (X-AXIS), WHEN CHANGING THE LEVEL OF IDIOSYNCRATIC RISK IN THE GERMAN BENCHMARK ECONOMY

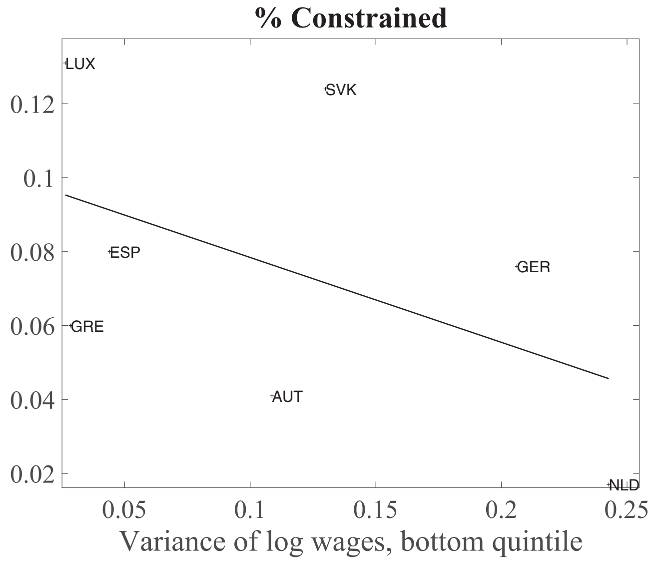
drop in response to fiscal consolidations should be smaller in countries with higher household debt because they have more constrained agents. We expand the Blanchard and Leigh (2014) regression with an interaction term between household debt and the planned fiscal consolidation and find exactly this: household debt diminishes the recessive effects of fiscal consolidation. The larger the household debt, the smaller the forecast error.

Second, to test how the labor supply elasticity to a fiscal consolidation shock depends on wealth, we use the PSID data set and the fiscal consolidation episodes identified by Alesina et al. (2015a) and used in Subsection 3.2. We establish that poor agents have a lower labor supply elasticity to public debt-altering fiscal shocks than their wealthier counterparts.

8.1. *Household Debt.* Blanchard and Leigh (2014) test whether precrisis household debt was one of the dimensions the IMF did not properly take into consideration when forecasting GDP growth rates. Similarly to all the other variables they examine, they find that debt does not affect the forecast error. However, our mechanism suggests that debt should have affected the recessive impacts of fiscal consolidation programs. Decreasing risk induces less precautionary savings, which results in higher household debt and consequently in a higher share of credit-constrained agents, as can be seen in Figure 12. We find some supporting empirical evidence detailed in Figure 13, which shows that higher income risk—proxied by the volatility of wages in the bottom 20% of the income distribution—is associated with a lower percentage of credit-constrained households. Higher household debt should, according to our model, translate into smaller multipliers.

To test whether household debt helps to explain the impacts of fiscal consolidation programs—besides extending Equation (1) with precrisis household debt, as already done by Blanchard and Leigh (2014)—we also include an interaction term between planned fiscal consolidation and precrisis household debt. The equation that we estimate is

$$\begin{aligned}\Delta Y_{i,t:t+1} - \hat{E}\{\Delta Y_{i,t:t+1}|\Omega_t\} &= \alpha + \beta \hat{E}\{F_{i,t:t+1|t}|\Omega_t\} + \gamma HD_{i,t-1} \\ &\quad + \iota((\hat{E}\{F_{i,t:t+1|t}|\Omega_t\})(HD_{i,t-1} - \mu_{HD})) + \epsilon_{i,t:t+1},\end{aligned} \tag{15}$$



NOTES: Correlation coefficient of −0.48 and p-value of 0.27. Wage data come from Luxembourg Income Study survey year 2010, which corresponds most closely to the survey responses of HFCS 1st wave. Constrained households are households who answered affirmatively to survey questions regarding either being rejected when seeking credit, not receiving full amount requested, or forgoing requesting credit due to expectation of being rejected.

FIGURE 13

THE RELATIONSHIP BETWEEN WAGE VOLATILITY OF THE LOWEST 20% OF EARNERS AND THE SHARE OF HOUSEHOLDS WHO FELT CREDIT CONSTRAINED

TABLE 6

BLANCHARD AND LEIGH (2013) REGRESSIONS AUGMENTED WITH HOUSEHOLD DEBT

Coefficients	(1) Blanchard–Leigh	(2) Blanchard–Leigh Precrisis Household Debt	(3) Precrisis Household Debt
β	−1.095***	−1.086***	−1.389***
	(0.255)	(0.262)	(0.117)
γ		−0.001	−0.004
		(0.006)	(0.003)
ι			0.010***
			(0.001)
Constant	0.775*	0.887	1.422***
	(0.383)	(0.699)	(0.420)
Observations	26	25	25
R^2	0.496	0.489	0.690

NOTES: The table displays the results from augmenting the regression in Blanchard and Leigh (2013) with household debt and an interaction term between household debt and planned fiscal consolidation. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. Robust standard errors in parentheses.

where $HD_{i,t-1}$ is precrisis household debt in country i, measured as total financial liabilities in percent of household disposable income (see data description in the Appendix). We use precrisis household debt so that it is exogenous to the fiscal shocks and to the output variation. Once again, we reparametrize the interaction term.

The results in Table 6 are consistent with our mechanism. The interaction term is positive and statistically significant. Moreover, the R^2 is substantially higher than in the specification without the interaction term, and the coefficient associated with the planned consolidation is more negative and statistically different from the specification without the interaction. This suggests that, during the consolidations in the European countries in 2010 and 2011, higher precrisis household debt contributed to diminishing the recessive effects of fiscal consolidation programs, just as our mechanism suggests. Increasing precrisis household debt by one standard deviation decreases the recessive impacts of fiscal consolidation by 52%.[45,46]

8.2. *Wealth and Labor Supply Responses to Fiscal Consolidation Shocks.* In the previous section, we have provided empirical evidence showing that the recessive impact of fiscal consolidation is decreasing in the percentage of credit-constrained agents, just as our mechanism suggests. The only part of our mechanism still missing validation is how the labor supply elasticities depend on wealth. Remember that, in our model, credit-constrained households are the least responsive to the shock.

To test whether the poorest agents are more inelastic to fiscal consolidation shocks, we use the PSID data set, which includes both wealth and hours worked, and combine it with IMF consolidation shocks from Alesina et al. (2015a) that span from 1987 to 2014. As PSID wealth data are only collected every two years, we sum the IMF shocks over two years and use that sum as the consolidation shocks.[47] Then, we proceed by regressing the first difference of the log of hours worked, $\Delta \ln h_{jt}$, for household j in year t on the sum of the unanticipated shocks over the previous two years, $\sum_{i=1}^{2} e^{u}_{t-i}$, and announced ones, $\sum_{i=1}^{2} e^{a}_{t-i}$, on the household wealth in the previous year a_{jt-1} and on interaction terms between wealth and the two fiscal consolidation shocks variables.[48] In addition, we control for household fixed effects.

$$\begin{aligned} \Delta \ln h_{jt} = {} & c + \beta_1 \sum_{i=1}^{2} e^{a}_{t-1} + \beta_2 \sum_{i=1}^{2} e^{u}_{t-1} + \beta_3 a_{jt-1} + \beta_4 a_{jt-1} \sum_{i=1}^{2} e^{a}_{t-1} \\ & + \beta_5 a_{jt-1} \sum_{i=1}^{2} e^{u}_{t-1} + \alpha_j + \epsilon_{jt}. \end{aligned} \tag{16}$$

As we established in Subsection 3.2, we could only find a statistically significant effect of the unanticipated component of consolidation shocks and its interaction with income inequality. Our finding for wealth inequality is similar. Thus, the coefficient of interest here is β_5, which captures the response of hours worked to an unanticipated fiscal shock depending on the asset position of the household. For the results to be in accordance with the model predictions, β_5 would have to be negative. This would translate into agents with more wealth dropping hours worked relatively more than the poor in response to an unanticipated consolidation shock. According to the theory, this happens because the wealth poor have a lower intertemporal elasticity of substitution.

The regression results are presented in Table 7. In Column (1), the β_2-coefficient shows that, in accordance with the model results, an unanticipated fiscal consolidation shock leads to a drop in labor supply. In Column (2), we include the interaction terms between liquid wealth

[45] In the Appendix, we also test for correlation with other measures that are closely related to our mechanism such as the percentage of credit-constrained households or household credit excluding mortgages as percentage of gross income, from the 2010 survey of Household Finance and Consumption Survey. The limitation of this data is that we only have 12 and 13 observations. The regression results for these two variables can be found in Table A14. Despite the coefficients not being statistically significant, the sign is in accordance with the model prediction.

[46] To check whether the result is robust to controlling for income inequality we run the same regression while including the different income inequality measures used in Subsection 3.1 and the interaction term with fiscal consolidations. Table A15 shows the household debt results to be robust to controlling for income inequality. We also test if the results are robust to including only the set of countries used to calibrate the model. The results in Table A16 shows that the results still go through if we only include this sample of countries.

[47] Only after 1999 did the PSID surveys begin being conducted every two years. Before that, wealth data were collected only every five years. The PSID surveys we include in this exercise are from the following years: 1989, 1994, 1999, 2001, 2003, 2005, 2007, 2009, 2011, 2013, and 2015.

[48] The definition of wealth used here is total wealth net of debt excluding home equity. In Table A17, we test the same regression with total wealth net of debt including home equity. The results still go through for this alternative measure of wealth. In the two cases, both wealth variables and change in hours worked are winsorized at the 1st and 99th percentile to control for outliers.

TABLE 7
LABOR SUPPLY RESPONSE, WEALTH, AND FISCAL CONSOLIDATION SHOCKS

VARIABLES	(1) $\Delta \ln h$	(2) $\Delta \ln h$	(3) $\Delta \ln h$
β_1	0.130***	0.130***	
	(0.023)	(0.023)	
β_2	−0.041	−0.043	
	(0.046)	(0.046)	
β_3	−0.072***	−0.060***	−0.019
	(0.013)	(0.014)	(0.014)
β_4		−0.022	−0.027
		(0.031)	(0.031)
β_5		−0.129**	−0.135***
		(0.050)	(0.051)
Constant	−0.165***	−0.166***	−0.148***
	(0.009)	(0.009)	(0.007)
Observations	72,969	72,969	72,969
R^2	0.123	0.123	0.126
Household FE	Yes	Yes	Yes
Year FE	No	No	Yes

NOTES: The table displays the results from estimating the regression in Equation (16). Robust standard errors in parentheses. ***$p < 0.01$, **$p < 0.05$, and *$p < 0.1$.

and both fiscal consolidation shocks. As expected, β_5 is negative and statistically significant, in accordance with the model predictions, translating into wealthier agents responding more to an unanticipated consolidation shock than wealth poor agents. Note as well that β_4 is not statistically significant, in line with the results in Subsection 3.2 not finding a significant relation between the propagation of anticipated consolidations and inequality. In the last column, (3), we control for year fixed effects and exclude consolidation shocks in levels. The results are robust to this alternative specification.[49]

The mechanism, for both the spending-based and tax-based consolidations, hinges on the intertemporal elasticity of substitution being higher for wealthier agents. Wealthier agents are more forward looking and should respond more to a future positive wage shock (the case of spending-based consolidation). Likewise they should respond more to a current cut in the wage rate (the case of tax-based consolidation). This is exactly what Domeij and Floden (2006) find. Using data from the PSID, they show that the response to a current wage rate shock is increasing by wealth quantile with the poorest agents responding the least and the wealthiest the most. Similarly, using an indicator of liquidity constraints, they find that constrained agents respond the least to a current wage rate shock.

The results of Domeij and Floden (2006) are very much in line with our above results where we combine PSID data with data on identified fiscal consolidation shocks to show that the labor supply of wealth poor agents responds the least to fiscal consolidations.

8.2.1. *Income risk and liquidity-constrained households.* Sections 8.1 and 8.2 establish the relationship between credit constraints and labor supply responses to fiscal consolidation shocks. Next, one may ask if more income risk leads to fewer credit-constrained households. This is a standard finding in the macroeconomic literature with incomplete markets, however, is there empirical evidence for this? Using the survey response about credit-constrained households in the HFCS Wave 1, we obtain data on credit-constrained households.

[49] Due to data limitations, we could not perform this exercise for European countries. To work around this, we use microdata from the labor force surveys (LFS) from EUROSTAT, and using age as proxy for wealth, test how the agents' age affects the labor supply response to fiscal shocks. We also test how the average hours worked per agent across countries responds to the fiscal consolidation and how this response is affected by income inequality in the economy. Both results are in accordance with the model mechanism, see the Appendix.

Constrained households are households who answered affirmatively to survey questions regarding either being rejected when seeking credit, not receiving full amount requested, or forgoing requesting credit due to expectation of being rejected. For years that match up with the HFCS survey (i.e., 2010), we also obtain data on wage volatility (measured as the variance of wages) for the bottom 20% of wages from LIS. This measure, the volatility of wages for low earners, should be closely related to precautionary savings. Although the sample size is small (seven countries: AT, DE, GR, LU, NL, SK, and ES), the correlation is negative (−0.5), see Figure 13 in Subsection 8.1, providing further evidence of the desired relationship.[50]

9. REPLICATION OF BLANCHARD AND LEIGH (2013) WITH MODEL DATA

In this section, we reproduce the empirical exercise in Subsection 3.1 using simulated data from our model. To replicate the empirical exercise, we need to create the forecast for the output response to the fiscal shock. In Subsection 3.1, we document that income inequality was one variable that the IMF failed to properly take into consideration when predicting the effects of fiscal consolidation programs (by augmenting the exercise in Blanchard and Leigh (2013)). In Sections 6 and 7, we establish that only the stochastic component of the income process can explain this relationship between inequality and the effects of the fiscal consolidation. Therefore, we identify our "forecasts" as consisting of the model output response to the fiscal consolidation shock when shutting down the stochastic component of the income process. In other words, we assume that the IMF had a model that was similar to ours, except that it did not have idiosyncratic income risk.[51] We then assume that the actual output response is given by our benchmark model, which properly models income risk.

For each of the 13 economies considered here, we calibrate the consolidation accordingly, matching them with the data on planned fiscal consolidations used in Subsection 3.1. The data are reported in Table A19.

The results from estimating Equation (2) for spending- and tax-based consolidation, using data from the model simulations, are presented in Tables 8 and A24, respectively.[52] Note that, as in Table 1, the coefficients for the interaction between the different measures of income inequality and both types of consolidations indicate that the effects of fiscal consolidation are amplified by income inequality.[53] The results for both the spending-based and the tax-based consolidations are statistically significant. Regardless of only having 13 observations and not matching any measure of labor income inequality to calibrate the model, the results are remarkably similar to the empirical ones presented in Subsection 3.1

Just as in the empirical exercise, the implications of abstracting from inequality are economically meaningful. A one-standard deviation increase in income inequality leads to an underestimation of the multiplier by 35% and 52%, for spending- and tax-based consolidation, respectively.

10. CONCLUSION

In this article, we use three independent data sources and three different empirical approaches to document a positive relationship between income inequality and the recessive impacts of fiscal consolidation programs. Income inequality is an important factor to account for when quantifying the impacts of fiscal consolidation.

[50] We also compute the correlation between the income inequality measures used in the empirical exercise in Section 9 and the percentage of credit-constrained agents. The correlation coefficients are negative across all income inequality measures, see Table A18.

[51] By assumption, the IMF model had some income inequality modeled as variation in permanent abilities but it was similar for all countries.

[52] Both spending- and tax-based consolidations are presented in absolute values, so an increase in both variables translates into a stronger consolidation.

[53] The inequality measures reflect total income inequality, as in Subsection 3.1.

TABLE 8
BLANCHARD AND LEIGH (2013) REGRESSIONS WITH MODEL-GENERATED FORECAST ERRORS FOR G-CONSOLIDATIONS

Coefficients	(1) Blanchard–Leigh	(2) Y25/Y75	(3) Y20/Y80	(4) Y10/Y90	(5) Y5/Y95	(6) Y2/Y98	(7) Income Gini
G consolidation	−0.610***	−0.634***	−0.654***	−0.747***	−0.879***	−0.844***	−0.623***
	(0.154)	(0.097)	(0.095)	(0.099)	(0.113)	(0.129)	(0.123)
Inequality		−0.064	−0.028	0.010	0.021	0.013	−0.097
		(0.155)	(0.110)	(0.049)	(0.027)	(0.019)	(0.081)
Interaction		−0.073	−0.072*	−0.068***	−0.063***	−0.040***	0.002
		(0.053)	(0.038)	(0.019)	(0.017)	(0.011)	(0.027)
Constant	−0.652*	−0.321	−0.447	−0.640	−0.772	−0.723	2.283
	(0.348)	(0.803)	(0.744)	(0.647)	(0.613)	(0.576)	(2.208)
Observations	13	13	13	13	13	13	13
R^2	0.370	0.444	0.444	0.450	0.456	0.442	0.510

NOTES: The table displays the results from running the regression in Equation (2), using model-generated forecast errors for spending-based consolidations. Y25/Y75, Y20/Y80, Y10/Y90, Y5/Y95, and Y2/Y98 represent the share of income of the top 25%, 20%, 10%, 5%, and 2% divided by the share of the bottom 25%, 20%, 10%, 5%, and 2%. Robust standard errors in parentheses. ***$p < 0.01$, **$p < 0.05$, and *$p < 0.1$.

To explain the amplification effect that income inequality has on the recessive impacts of fiscal consolidation programs, we develop a life cycle, overlapping generations economy with uninsurable labor market risk. We calibrate the model to data from European economies and study the effects of fiscal consolidation programs, financed through both austerity and taxation. We find that if cross-country differences in inequality are due to income risk, then the model can explain the relationship between inequality and the effect of consolidation on output. Differences in initial conditions, modeled as permanent ability and the life cycle profile of wages, cannot account for the cross-country variation in the impacts of fiscal consolidation we observe in the data.

The relationship between risk and the impact of consolidation arises because, in countries with higher income risk, agents will have higher savings due to precautionary motives and thus there will be a smaller share of credit-constrained and low-wealth agents. These agents have a lower intertemporal elasticity of substitution. Wealthier agents are more forward looking and respond more to a future positive income shock (the case of spending-based consolidation). Similarly, they respond more to a current cut in the wage rate (the case of tax-based consolidation).

A decrease in government debt leads to an increase in productive capital in the economy. The marginal product of labor (wages) in future time periods increases and this is equivalent to a permanent positive income shock, causing labor supply and output to fall in the short run. The response is, however, smaller in economies with more credit-constrained agents. A tax-based consolidation also decreases current wages. Wealthy agents will be able to consume their savings and reduce current labor supply, whereas poor agents will have to work hard to reduce the fall in current consumption. Also for tax-based consolidations, the response is smaller in economies with more credit-constrained agents.

To show that the mechanism we propose is consistent with the data, we conduct two empirical exercises. First, making use of the positive correlation between household debt and credit-constrained agents, we establish that countries with higher household debt experience a smaller output drop during a fiscal consolidation. This is just as our mechanism suggests. Second, we show that the labor supply response to fiscal consolidation shocks is increasing in wealth, with poor agents responding the least to these shocks, just as our mechanism predicts.

There are still many open questions regarding the fiscal policy transmission mechanisms. Nonetheless, we present evidence showing that income inequality is an important determinant of the impacts of fiscal consolidation programs.

APPENDIX A.

A.1 *Technology.* There is a representative firm, producing output with a Cobb–Douglas production function:

$$Y_t(K_t, L_t) = K_t^{\alpha} L_t^{1-\alpha}, \tag{A.1}$$

where K_t is the capital input and L_t the labor input in efficiency units. The evolution of capital is given by

$$K_{t+1} = (1-\delta)K_t + I_t, \tag{A.2}$$

where I_t is gross investment and δ the capital depreciation rate. Each period, the firm hires labor and capital to maximize its profits:

$$\Pi_t = Y_t - w_t L_t - (r_t + \delta)K_t. \tag{A.3}$$

In a competitive equilibrium, the factor prices will be equal to their marginal products given by

$$w_t = \partial Y_t/\partial L_t = (1-\alpha)\left(\frac{K_t}{L_t}\right)^{\alpha}, \tag{A.4}$$

$$r_t = \partial Y_t/\partial K_t - \delta = \alpha\left(\frac{L_t}{K_t}\right)^{1-\alpha} - \delta. \tag{A.5}$$

A.2 *Tax Function.* Given the tax function[54]

$$ya = \theta_0 y^{1-\theta_1},$$

which we employ, the after tax income is defined as

$$ya = (1-\tau_l(y))y,$$

and thus

$$\theta_0 y^{1-\theta_1} = (1-\tau_l(y))y,$$

and thus

$$\begin{aligned} 1-\tau_l(y) &= \theta_0 y^{-\theta_1} \\ \tau_l(y) &= 1-\theta_0 y^{-\theta_1} \\ T_l(y) &= \tau_l(y)y = y - \theta_0 y^{1-\theta_1} \\ T_l'(y) &= 1-(1-\theta_1)\theta_0 y^{-\theta_1}. \end{aligned}$$

Thus the tax wedge for any two incomes (y_1, y_2) is given by

$$1 - \frac{1-\tau_l(y_2)}{1-\tau_l(y_1)} = 1 - \left(\frac{y_2}{y_1}\right)^{-\theta_1}, \tag{A.6}$$

and therefore independent of the scaling parameter θ_0. Thus by construction one can raise average taxes by lowering θ_0 and not change the progressivity of the tax code, since (as long as tax progressivity is defined by the tax wedges) the progressivity of the tax code [55] is uniquely determined by the parameter θ_1.

A.3 *Definition of a Stationary Recursive Competitive Equilibrium.* Let the measure of households with the corresponding characteristics be given by $\Phi(k, \beta, a, u, j)$. The stationary recursive competitive equilibrium is defined by:

[54] This appendix is borrowed from Holter et al. (2019).

[55] Note that

$$1-\tau_l(y) = \frac{1-T_l'(y)}{1-\theta_1} > 1-T_l'(y).$$

and thus as long as $\theta_1 \in (0,1)$, we have that

$$T_l'(y) > \tau_l(y)$$

and thus marginal tax rates are higher than average tax rates for all income levels.

(1) Given the factor prices and the initial conditions the consumers' optimization problem is solved by the value function $V(k, \beta, a, u, j)$ and the policy functions, $c(k, \beta, a, u, j)$, $k'(k, \beta, a, u, j)$, and $n(k, \beta, a, u, j)$.

(2) Markets clear:

$$K + B = \int k d\Phi$$

$$L = \int (n(k, \beta, a, u, j)) d\Phi$$

$$\int c d\Phi + \delta K + G = K^{\alpha} L^{1-\alpha}.$$

(3) The factor prices satisfy

$$w = (1-\alpha)\left(\frac{K}{L}\right)^{\alpha},$$

$$r = \alpha\left(\frac{K}{L}\right)^{\alpha-1} - \delta.$$

(4) The government budget balances:

$$g \int d\Phi + G + rB = \int \left(\tau_k r(k+\Gamma) + \tau_c c + n\tau_l\left(\frac{nw(a, u, j)}{1+\tilde{\tau}_{ss}}\right)\right) d\Phi.$$

(5) The social security system balances

$$\Psi \int_{j \geq 65} d\Phi = \frac{\tilde{\tau}_{ss} + \tau_{ss}}{1+\tilde{\tau}_{ss}} \left(\int_{j<65} nwd\Phi\right).$$

(6) The assets of the dead are uniformly distributed among the living:

$$\Gamma \int \omega(j) d\Phi = \int (1-\omega(j)) k d\Phi.$$

A.4 *Definition of a Transition Equilibrium After the Unanticipated Fiscal Consolidation Shock.* We define a recursive competitive equilibrium along the transition between steady states as follows:

Given the initial capital stock, the initial distribution of households and initial taxes, respectively, K_0, Φ_0, and $\{\tau_l, \tau_c, \tau_k, \tau_{ss}, \tilde{\tau}_{ss}\}_{t=1}^{t=\infty}$, a competitive equilibrium is a sequence of individual functions for the household, $\{V_t, c_t, k'_t, n_t\}_{t=1}^{t=\infty}$, of production plans for the firm, $\{K_t, L_t\}_{t=1}^{t=\infty}$, factor prices, $\{r_t, w_t\}_{t=1}^{t=\infty}$, government transfers $\{g_t, \Psi_t, G_t\}_{t=1}^{t=\infty}$, government debt, $\{B_t\}_{t=1}^{t=\infty}$, inheritance from the dead, $\{\Gamma_t\}_{t=1}^{t=\infty}$, and of measures $\{\Phi_t\}_{t=1}^{t=\infty}$, such that for all t:

(1) Given the factor prices and the initial conditions the consumers' optimization problem is solved by the value function $V(k, \beta, a, u, j)$ and the policy functions, $c(k, \beta, a, u, j)$, $k'(k, \beta, a, u, j)$, and $n(k, \beta, a, u, j)$.

(2) Markets clear:

$$K_{t+1} + B_t = \int k_t d\Phi_t$$

$$L_t = \int (n_t(k_t, \beta, a, u, j))d\Phi_t$$

$$\int c_t d\Phi_t + K_{t+1} + G_t = (1-\delta)K_t + K_t^{\alpha} L_t^{1-\alpha}.$$

(3) The factor prices satisfy

$$w_t = (1-\alpha)\left(\frac{K_t}{L_t}\right)^{\alpha},$$

$$r_t = \alpha\left(\frac{K_t}{L_t}\right)^{\alpha-1} - \delta.$$

(4) The government budget balances:

$$g_t \int d\Phi_t + G_t + r_t B_t = \int \left(\tau_k r_t(k_t + \Gamma_t) + \tau_c c_t + n_t \tau_l \left(\frac{n_t w_t(a,u,j)}{1+\tilde{\tau}_{ss}}\right)\right) d\Phi_t + (B_{t+1} - B_t).$$

(5) The social security system balances

$$\Psi_t \int_{j\geq 65} d\Phi_t = \frac{\tilde{\tau}_{ss} + \tau_{ss}}{1+\tilde{\tau}_{ss}} \left(\int_{j<65} n_t w_t d\Phi_t\right).$$

(6) The assets of the dead are uniformly distributed among the living:

$$\Gamma_t \int \omega(j) d\Phi_t = \int (1-\omega(j)) k_t d\Phi_t.$$

(7) Aggregate law of motion:

$$\Phi_{t+1} = \Upsilon_t(\Phi_t).$$

A.5 *Definition of the Fiscal Multiplier in the Context of a Fiscal Consolidation Shock.*

$$impact\, multiplier = \frac{\Delta Y_0}{\Delta X_0}, \tag{A.7}$$

where ΔY_0 is the change in output from period 0 to period 1 and ΔX_0 is the change in either government spending or labor tax revenues from period 0 to period 1. The cumulative multiplier at time T is defined as

$$cumulative\, multiplier(T) = \frac{\sum_{t=0}^{t=T} \left(\Pi_{s=0}^{s=t} \frac{1}{(1+r_s)}\right) \Delta Y_t}{\sum_{t=0}^{t=T} \left(\Pi_{s=0}^{s=t} \frac{1}{(1+r_s)}\right) \Delta X_t}, \tag{A.8}$$

where ΔY_t is the change in output from period 0 to period t and ΔX_0 is the change in either government spending or labor tax revenues from period 0 to period t.

A.6 *Description of Data Used in Sections 3, 8, and 9.* The data series used in Subsections 3.1 and 3.2 for the inequality measures are from the EU-SILC. The EU-SILC is a survey aiming at collecting cross-sectional and longitudinal microdata on income, poverty, social exclusion, and living-conditions. Data collected is based on a nationally representative probability

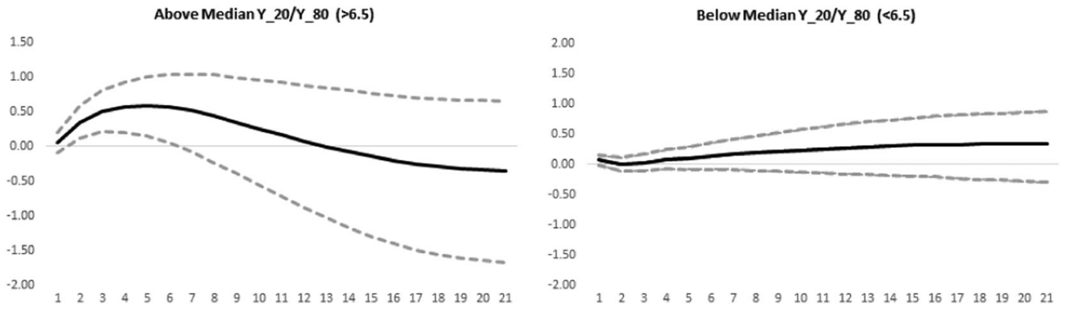


FIGURE A1

CUMULATIVE OUTPUT MULTIPLIER, AS DEFINED IN (A.11), TO A GOVERNMENT CONSUMPTION SHOCK (90% ERROR BANDS IN GRAY)

sample of the population residing in private households within the country. Cross-sectional data series used is gross income – total monetary and nonmonetary income received by the household before deduction of taxes.

The growth forecast error and planned fiscal consolidation series are taken from Blanchard and Leigh (2014), who use data from the IMF's WEO database. The forecasts used were made for European economies in early 2010. The growth forecast errors are the difference between actual cumulative growth in 2010–2011 and the IMF forecast prepared for the April 2010 WEO. The planned fiscal consolidation is the IMF forecast of the cumulative changes of structural fiscal balance as percent of potential GDP, also prepared for the April 2010 WEO. The household debt variable used in Subsection 8.1 also comes from Blanchard and Leigh (2014), who take it from the data set of the April 2012 WEO chapter on household debt. Household debt consists of total financial liabilities in percent of household disposable income.

The data series used in Section A.8 are taken from Ilzetzki et al. (2013). The data series consist of quarterly observations (not interpolated) on real government consumption, GDP, the ratio of current account to GDP, and the real effective exchange rate for 44 countries, roughly balanced between developed and developing economies (see Table A1 for the list of included countries). Nominal series are deflated using a GDP deflator when available (and CPI when not). Consumption, GDP, and exchange-rate variables are transformed by taking natural logarithms. These series are de-seasonalized and analyzed as deviations from their quadratic trend given they exhibit strong seasonality and are nonstationary. Data in Table A1 come from the World Bank's World Development Indicators for the years of 2009 for Botswana and Malaysia, 2010 for Australia, Canada, and Israel, 2011 for Germany and South Africa, 2012 for Belgium, Bulgaria, Croatia, Czech Republic, Denmark, Estonia, Finland, France, Greece, Hungary, Iceland, Ireland, Italy, Latvia, Lithuania, Mexico, Netherlands, Norway, Portugal, Slovakia, Slovenia, Spain, Sweden, and United Kingdom, and 2013 for all the other countries.

A.7 *Eurosystem Household Finance and Consumption Survey—Summary Wealth Statistics.* Table A2 presents the cumulative wealth distributions for the countries in the Eurosystem Household Finance and Consumption Survey. We include four additional countries' wealth distributions from other surveys when available, for example, the LWS' compilation of various household wealth surveys.

A.8 *SVAR.* In this subsection, we provide additional evidence on the link between income inequality and the recessive impacts of fiscal contractions, using a larger data set containing 44 countries, see data description in Section A.6. We use the data and methodology from Ilzetzki et al. (2013) to run VARs for two different groups of countries pooled by their

TABLE A1
INCOME INEQUALITY MEASURES FOR 44 SELECTED COUNTRIES

Country	Income Gini	Y20/Y80	Y10/Y90
Argentina	0.42	9.8	19.1
Australia	0.35	5.9	10.2
Belgium	0.28	4.2	6.7
Botswana	0.61	23.1	45.0
Brazil	0.53	17.4	41.8
Bulgaria	0.36	6.9	13.7
Canada	0.34	5.8	9.5
Chile	0.51	12.3	24.4
Colombia	0.54	17.1	38.1
Croatia	0.33	5.7	9.6
Czech Republic	0.26	3.8	5.7
Denmark	0.29	4.4	8.4
Ecuador	0.47	11.5	22.8
El Salvador	0.43	9.1	16.4
Estonia	0.33	5.7	10.1
Finland	0.28	3.9	5.7
France	0.33	5.3	8.6
Germany	0.30	4.6	7.0
Greece	0.37	7.6	15.7
Hungary	0.31	4.9	8.0
Iceland	0.27	4.0	6.0
Ireland	0.33	5.3	8.3
Israel	0.43	10.3	18.4
Italy	0.35	6.7	13.8
Latvia	0.36	6.7	12.1
Lithuania	0.35	6.5	11.7
Malaysia	0.46	11.2	19.2
Mexico	0.48	11.0	20.5
Netherlands	0.28	4.2	6.6
Norway	0.26	3.8	5.8
Peru	0.45	11.3	22.3
Poland	0.33	5.2	7.8
Portugal	0.36	6.6	12.6
Romania	0.28	4.1	6.0
Slovakia	0.26	4.1	6.6
Slovenia	0.26	3.7	5.7
South Africa	0.63	27.6	57.0
Spain	0.36	7.2	15.2
Sweden	0.27	4.2	6.7
Thailand	0.38	6.5	10.1
Turkey	0.40	8.0	13.9
United Kingdom	0.33	5.4	8.5
United States	0.41	9.1	17.8
Uruguay	0.42	9.3	16.3
Sample median	0.35	6.5	10.9

position above or below the median income inequality in the sample. We use three different measures of inequality: (i) the income share of the top 20% divided by the share of the bottom 20%; (ii) the income share of the top 10% divided by the income share of the bottom 10%; (iii) the income Gini coefficient. We find that the results are consistent across the three different metrics of income inequality. For countries with the income inequality metric above the median, the recessive impacts of decreases in government consumption expenditures are stronger and statistically different from the impacts for the group of countries with income inequality metrics below the median.

TABLE A2
CUMULATIVE DISTRIBUTION OF NET WEALTH

	25%	50%	75%
HFCS sample[a]			
Austria	−1.0	2.2	18.6
France	0.1	5.4	26.2
Germany	−0.4	2.7	17.9
Greece	1.1	12.5	36.7
Italy	0.9	10.2	32.4
Netherlands	−2.5	5.0	30.3
Portugal	0.6	8.2	26.6
Slovakia	5.5	20.7	45.0
Spain	1.7	12.9	34.2
Other sources			
Czech Republic[b]	0.4	6.1	22.1
Iceland[b]	0.5	7.7	27.6
Sweden[c]	−9.9	−7.8	11.5
United Kingdom[c]	−0.7	5.4	27.0

[a]Cumulative distribution of net wealth (survey variable designation: *DN3001*) for a selection of countries from the first wave of the ECB's HFCS. Units: percent of total wealth.
[b]Sourced from the Stierli et al. (2014). We use 2009 data provided by the authors.
[c]Sourced from Luxembourg Wealth Study's entry nearest to 2008 for each respective country (survey variable designation: *nw1*).

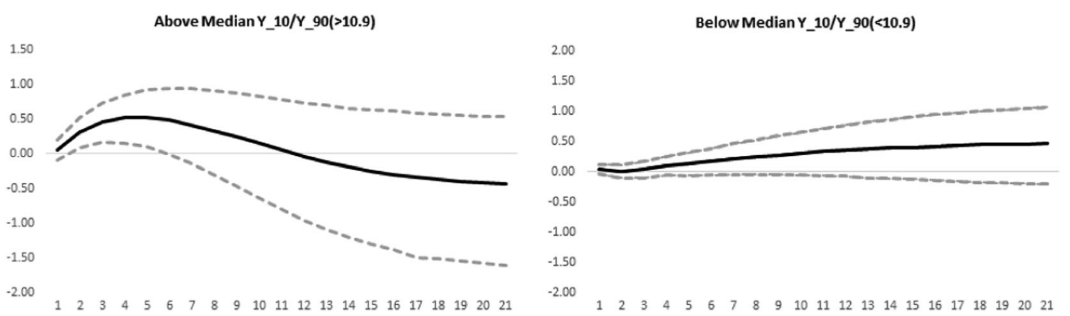


FIGURE A2

CUMULATIVE OUTPUT MULTIPLIER, AS DEFINED IN (A.11), TO A GOVERNMENT CONSUMPTION SHOCK (90% ERROR BANDS IN GRAY)

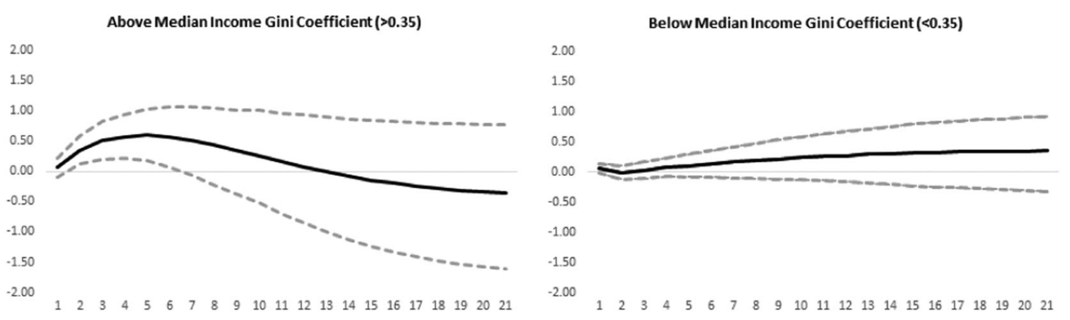


FIGURE A3

CUMULATIVE OUTPUT MULTIPLIER, AS DEFINED IN (A.11), TO A GOVERNMENT CONSUMPTION SHOCK (90% ERROR BANDS IN GRAY)

TABLE A3

MONTE CARLO EXPERIMENT WITH $\rho = 0.38$ TO ASSESS NICKEL'S BIAS. AT $T = 30$, THE NEGATIVE BIAS IS AROUND 12%

T	Average Absolute Bias	Average Variance
5	−0.3138	0.0987
10	−0.1504	0.0227
15	−0.0975	0.0096
20	−0.0720	0.0052
25	−0.0569	0.0033
30	−0.0474	0.0023
50	−0.0280	0.0008
100	−0.0138	0.0002
200	−0.0069	0.0001

The objective is to estimate the following system of equations

$$AY_{nt} = \sum_{k=1}^{K} C_k Y_{n,t-k} + u_{n,t}, \tag{A.9}$$

where Y_{nt} is a vector containing the endogenous variables for country n in quarter t. The variables considered are the same as in Ilzetzki et al. (2013): government consumption, output, current account in percentage of GDP, and the natural logarithm of the real effective exchange rate. C_k is a matrix of lag own and cross-effects of variables on their current observations. Given that A is not observable, we cannot estimate this regression directly. We need to pre-multiply everything by A^{-1} and, using OLS, we can recover the matrix $P = A^{-1}C_k$ and $e_{n,t} = A^{-1}u_{n,t}$. So we estimate the system

$$Y_{nt} = \sum_{k=1}^{K} A^{-1} C_k Y_{n,t-k} + A^{-1} u_{n,t}. \tag{A.10}$$

To be able to estimate the effects of fiscal consolidation, we need more assumptions on A so that we can identify the innovations by solving $e_{n,t} = A^{-1}u_{n,t}$. We use the same assumption used by Ilzetzki et al. (2013) and first introduced by Blanchard and Perotti (2002), to identify the responses of output to government consumption expenditures: government consumption cannot react to shocks in output within the same quarter. The plausibility of this assumption comes from the fact that the government's budget is typically set on a yearly basis and can only react to changes in output with a lag. For the ordering of the remaining variables, we also follow Ilzetzki et al. (2013) and let the current account follow output and the real exchange rate follow the current account. Given this, we can identify the impulse responses to a primitive shock in government spending. In Figures A1–A3, we plot the cumulative output multiplier to a government consumption shock, defined as

$$cummulative\, multiplier\, G(T) = \frac{\sum_{t=0}^{t=T} \left(\frac{1}{(1+r_m)}\right)^t \Delta Y_t}{\sum_{t=0}^{t=T} \left(\frac{1}{(1+r_m)}\right)^t \Delta G_t}, \tag{A.11}$$

where r_m is the median interest rate in the data sample. The output multipliers shown in Figures A1–A3 suggest that in countries with higher income inequality, contractions in government spending have a more recessive impact.

A.9 *Discussion of the Estimation of the Parameters in the Wage Function.* One should note that the estimation of the AR(1) process for u to obtain ρ depends crucially on the

TABLE A4
INCOME INEQUALITY AND THE LABOR SUPPLY RESPONSE TO FISCAL CONSOLIDATION SHOCKS

Coefficients	(1) Benchmark	(2) Y25/Y75	(3) Y20/Y80	(4) Y10/Y90	(5) Y5/Y95	(6) Y2/Y98	(7) Income Gini
β_1	−0.004***	−0.003**	−0.003*	−0.003**	−0.003***	−0.007***	−0.004**
	(0.001)	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)	(0.002)
β_2	−0.004***	−0.001	−0.002	−0.006***	−0.006***	−0.004**	0.000
	(0.001)	(0.002)	(0.002)	(0.002)	(0.001)	(0.001)	(0.003)
γ		0.319	0.191	0.100	0.026	0.023**	0.068
		(0.232)	(0.167)	(0.068)	(0.027)	(0.010)	(0.093)
ι_1		−0.116	−0.155	−0.134***	−0.045***	−0.014**	−0.029
		(0.123)	(0.103)	(0.045)	(0.015)	(0.006)	(0.040)
ι_2		−0.266	−0.161	0.091	0.044**	0.005	−0.114
		(0.206)	(0.171)	(0.070)	(0.020)	(0.006)	(0.068)
Constant	0.211***	0.188***	0.194***	0.196***	0.204***	0.204***	0.184***
	(0.001)	(0.016)	(0.014)	(0.010)	(0.007)	(0.003)	(0.036)
Observations	55	55	55	55	55	55	55
R^2	0.502	0.582	0.577	0.601	0.618	0.563	0.567
Number of countries	11	11	11	11	11	11	11

NOTES: The table displays the results from estimating the regression in Equation (A.15) on data from Alesina et al. (2015a) and measures of income inequality from the EU-SILC. Y25/Y75, Y20/Y80, Y10/Y90, Y5/Y95, and Y2/Y98 represent the share of income of the top 25%, 20%, 10%, 5%, and 2% divided by the share of the bottom 25%, 20%, 10%, 5%, and 2%. Standard errors in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, and $^{*}p < 0.1$.

TABLE A5
THE IMPACT OF AGE ON THE LABOR SUPPLY RESPONSE TO FISCAL CONSOLIDATION SHOCKS

Coefficients	(1) Benchmark	(2) Age Cohort Interaction
β_1	−0.205***	0.758***
	(0.028)	(0.158)
β_2	−0.356***	0.525***
	(0.035)	(0.149)
γ		0.297***
		(0.015)
ι_1		−0.154***
		(0.022)
ι_2		−0.136***
		(0.022)
Constant	39.990***	38.207***
	(0.001)	(0.124)
Observations	1,022,421	1,022,421
R^2	0.016	0.018
Number of countries	11	11

NOTES: The table displays the results from estimating the regression in Equation (A.16) on fiscal consolidation shocks from Alesina et al. (2015a) and worker hours and age from the EU LFS. The parameter of interest is ι_1, that is, the interaction of the unanticipated fiscal consolidation shock with age. It shows that, for the same size of an unanticipated fiscal consolidation shock, the older the cohort in which the agent is included, the steeper the drop in labor supply. Standard errors in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, and $^{*}p < 0.1$.

TABLE A6
BLANCHARD AND LEIGH (2013) REGRESSIONS AUGMENTED WITH ALTERNATIVE MEASURES OF INCOME INEQUALITY

VARIABLES	(1) Blanchard–Leigh	(2) Y5	(3) Y10	(4) Y20	(5) Y25
β	−1.095***	−0.866***	−0.944***	−0.870***	−0.853***
	(0.255)	(0.190)	(0.202)	(0.218)	(0.228)
γ		0.425	0.806	0.266	0.116
		(1.349)	(0.531)	(0.374)	(0.279)
ι		1.562**	0.264	0.265	0.276*
		(0.664)	(0.296)	(0.203)	(0.158)
Constant	0.775*	0.350	−1.871	−1.478	−0.521
	(0.383)	(1.776)	(1.842)	(3.258)	(3.227)
Observations	26	26	26	26	26
R^2	0.496	0.598	0.582	0.582	0.581

NOTES: The table displays the results from augmenting the regression in Blanchard and Leigh (2013) with different measures of income inequality and an interaction term between income inequality and planned fiscal consolidation. Y5, Y10, Y20, and Y25 represent the share of income of the bottom 5%, 10%, 20%, and 25%. Robust standard errors in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, and $^{*}p < 0.1$

specification of the wage function, the choice of the wage measurement (i.e., labor income or hourly wages) and the sample selection. These choices have all varied in the literature and given rise to quite a wide range of results.

Labor productivity in our article is given by: $w_i(j, a, u) = we^{\gamma_1 j+\gamma_2 j^2+\gamma_3 j^3+a+u}$. To estimate the unknown parameters in this wage equation, Brinca et al. (2016) first estimate a log(wage) equation using data on hourly wages from the PSID 1968–1997:

$$\log(w_{it}) = X\beta + \gamma_1 age_{it} + \gamma_2 age_{it}^2 + \gamma_3 age_{it}^3 + u_{it}, \tag{A.12}$$

where X is a vector with time dummies. This gives estimates of the parameters governing the age profile of wages, $\gamma_1, \gamma_2, \gamma_3$. The AR(1) process for ρ is estimated jointly with the variance of ability σ_a^2 using the residuals from the first regression. That is, to obtain the parameters, σ_ϵ,

TABLE A7
BLANCHARD AND LEIGH (2013) REGRESSIONS AUGMENTED WITH MEASURES OF DISPOSABLE INCOME INEQUALITY

Coefficients	(1) Blanchard–Leigh	(2) Y25/Y75	(3) Y20/Y80	(4) Y10/Y90	(5) Y5/Y95	(6) Y2/Y98	(7) Income Gini
β	−1.095***	−0.823***	−0.826***	−0.954***	−0.883***	−0.899***	−0.815***
	(0.255)	(0.224)	(0.218)	(0.206)	(0.191)	(0.230)	(0.256)
γ		−0.110	−0.204	−0.252	0.035	−0.002	−0.016
		(0.570)	(0.438)	(0.150)	(0.073)	(0.050)	(0.143)
ι		−0.667**	−0.439**	−0.052	−0.102***	−0.053*	−0.143*
		(0.292)	(0.209)	(0.081)	(0.035)	(0.026)	(0.082)
Constant	0.775*	1.328	1.791	2.699**	0.516	0.956	1.335
	(0.383)	(2.367)	(2.111)	(1.203)	(1.090)	(1.094)	(4.200)
Observations	26	26	26	26	26	26	26
R^2	0.496	0.605	0.606	0.595	0.598	0.562	0.588

NOTES: The table displays the results from augmenting the regression in Blanchard and Leigh (2013) with different measures of income inequality and an interaction term between income inequality and planned fiscal consolidation. Y25/Y75, Y20/Y80, Y10/Y90, Y5/Y95, and Y2/Y98 represent the share of income of the top 25%, 20%, 10%, 5%, and 2% divided by the share of the bottom 25%, 20%, 10%, 5%, and 2%. Robust standard errors in parentheses. ***$p < 0.01$, **$p < 0.05$, and *$p < 0.1$.

TABLE A8
BLANCHARD AND LEIGH (2013) REGRESSIONS WITH GDP AS THE DEPENDENT VARIABLE

Coefficients	Blanchard–Leigh	Y25/Y75	Y20/Y80	Y10/Y90	Y5/Y95	Y2/Y98	Income Gini
β	−1.556***	−1.116**	−1.078**	−0.901**	−0.901**	−1.026***	−1.696***
	(0.467)	(0.441)	(0.436)	(0.379)	(0.339)	(0.339)	(0.406)
γ		−0.402	−0.302	−0.170	−0.039	0.019	0.286
		(0.578)	(0.444)	(0.191)	(0.053)	(0.050)	(0.169)
ι		−0.405	−0.365	−0.229*	−0.116***	−0.098**	−0.000
		(0.388)	(0.304)	(0.113)	(0.036)	(0.035)	(0.115)
Constant	3.763***	6.545*	6.335*	6.264**	4.938***	3.612***	−6.990
	(0.576)	(3.760)	(3.493)	(2.578)	(1.302)	(1.057)	(6.402)
Observations	26	26	26	26	26	26	26
R^2	0.465	0.533	0.544	0.607	0.634	0.587	0.542

NOTES: The table displays the results from estimating the regression in (1) with GDP as the dependent variable. Y25/Y75, Y20/Y80, Y10/Y90, Y5/Y95 and Y2/Y98 represent the share of income of the top 25%, 20%, 10%, 5%, and 2% divided by the share of the bottom 25%, 20%, 10%, 5%, and 2%. Robust standard errors in parentheses. ***$p < 0.01$, **$p < 0.05$, and *$p < 0.1$.

TABLE A9
BLANCHARD AND LEIGH (2013) REGRESSIONS WITH LAGGED DEPENDENT VARIABLE

Coefficients	Blanchard−Leigh	Y25/Y75	Y20/Y80	Y10/Y90	Y5/Y95	Y2/Y98	Income Gini
β	−0.618***	−0.331*	−0.340**	−0.398**	−0.471***	−0.458***	−0.703***
	(0.143)	(0.166)	(0.163)	(0.162)	(0.165)	(0.161)	(0.146)
γ		−0.373	−0.274	−0.112	−0.018	−0.009	0.073
		(0.286)	(0.215)	(0.093)	(0.037)	(0.026)	(0.075)
ι		−0.200*	−0.168*	−0.072	−0.030	−0.032*	−0.017
		(0.109)	(0.085)	(0.043)	(0.019)	(0.017)	(0.052)
Y_{it-1}	1.203***	1.231***	1.203***	1.106***	1.101***	1.092***	1.098***
	(0.222)	(0.158)	(0.166)	(0.226)	(0.257)	(0.254)	(0.217)
Constant	−0.085	2.386	2.140	1.536	0.463	0.256	−2.734
	(0.264)	(1.695)	(1.535)	(1.151)	(0.777)	(0.580)	(2.729)
Observations	26	26	26	26	26	26	26
R^2	0.809	0.869	0.868	0.855	0.836	0.844	0.816

NOTES: The table displays the results from estimating the regression in (1) with GDP as the dependent variable. Y25/Y75, Y20/Y80, Y10/Y90, Y5/Y95, and Y2/Y98 represent the share of income of the top 25%, 20%, 10%, 5%, and 2% divided by the share of the bottom 25%, 20%, 10%, 5%, and 2%. Robust standard errors in parentheses. ***$p < 0.01$, **$p < 0.05$, and *$p < 0.1$.

TABLE A10
BLANCHARD AND LEIGH (2013) REGRESSIONS WITH VARIABLES WINSORIZED AT THE 95TH PERCENTILE

Coefficients	Blanchard–Leigh	Y25/Y75	Y20/Y80	Y10/Y90	Y5/Y95	Y2/Y98	Income Gini
β	−1.128***	−0.882***	−0.858***	−0.781**	−0.887***	−0.861***	−1.319***
	(0.349)	(0.292)	(0.298)	(0.290)	(0.271)	(0.271)	(0.381)
γ		−0.243	−0.167	−0.024	0.063	0.056	0.287**
		(0.411)	(0.319)	(0.143)	(0.049)	(0.039)	(0.122)
ι		−0.296	−0.290	−0.238***	−0.136***	−0.117***	−0.069
		(0.317)	(0.239)	(0.083)	(0.033)	(0.027)	(0.106)
Constant	0.700*	2.426	2.200	1.346	−0.228	−0.139	−9.975**
	(0.388)	(2.763)	(2.590)	(1.976)	(1.098)	(0.697)	(4.540)
Observations	26	26	26	26	26	26	26
R^2	0.408	0.474	0.492	0.581	0.594	0.606	0.549

NOTES: The table displays the results from estimating the regression in (1) with GDP as the dependent variable. Y25/Y75, Y20/Y80, Y10/Y90, Y5/Y95, and Y2/Y98 represent the share of income of the top 25%, 20%, 10%, 5%, and 2% divided by the share of the bottom 25%, 20%, 10%, 5%, and 2%. Robust standard errors in parentheses. ***$p < 0.01$, **$p < 0.05$, and *$p < 0.1$.

TABLE A11
REGRESSIONS ON DATA FROM ALESINA ET AL. (2015A) WITH LAGGED DEPENDENT VARIABLE

Coefficients	(1) Benchmark	(2) Y25/Y75	(3) Y20/Y80	(4) Y10/Y90	(5) Y5/Y95	(6) Y2/Y98	(7) Income Gini
β_1	−0.004	0.006	0.003	−0.005	−0.006	−0.005	0.010
	(0.006)	(0.007)	(0.007)	(0.006)	(0.006)	(0.007)	(0.007)
β_2	−0.003	−0.003	−0.002	−0.001	−0.003	0.001	−0.001
	(0.005)	(0.007)	(0.007)	(0.007)	(0.006)	(0.006)	(0.007)
γ		−2.214**	−1.245	0.010	0.040	0.008	−1.081***
		(1.001)	(0.759)	(0.349)	(0.136)	(0.049)	(0.378)
ι_1		−1.517**	−0.986*	0.092	0.068	−0.004	−0.557***
		(0.603)	(0.511)	(0.234)	(0.077)	(0.030)	(0.195)
ι_2		−0.471	−0.284	−0.100	−0.015	0.024	−0.152
		(0.639)	(0.515)	(0.246)	(0.091)	(0.026)	(0.174)
Y_{it-1}	−0.075	−0.132	−0.116	−0.074	−0.076	−0.082	−0.144
	(0.112)	(0.112)	(0.114)	(0.116)	(0.114)	(0.116)	(0.108)
Constant	0.017***	0.169**	0.121*	0.015	0.006	0.014	0.430***
	(0.006)	(0.069)	(0.063)	(0.051)	(0.035)	(0.015)	(0.145)
Observations	84	84	84	84	84	84	84
R^2	0.014	0.150	0.100	0.018	0.036	0.028	0.201
Number of countries	12	12	12	12	12	12	12

NOTES: The table displays the results from estimating the regression in Equation (3) on data from Alesina et al. (2015a) and measures of income inequality from the EU-SILC. Y25/Y75, Y20/Y80, Y10/Y90, Y5/Y95, and Y2/Y98 represent the share of income of the top 25%, 20%, 10%, 5%, and 2% divided by the share of the bottom 25%, 20%, 10%, 5%, and 2%. Standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, and * $p < 0.1$.

ρ, and σ_a, they obtain the residuals u_{it} and use them to estimate the below equation by fixed effects estimation:

$$u_{it} = constant + a_i + \rho u_{it-1} + \epsilon_{it}. \tag{A.13}$$

The fixed effect or "ability" of an individual, a_i, is permanent throughout the life cycle. When specifying the residual wage equation like this, with a fixed effect and an AR(1) process in the lagged residual, one obtains relatively low estimates of ρ. In our benchmark specification, we obtain 0.34. This is a somewhat low value of ρ relative to much of the existing literature. We therefore decided to perform a battery of experiments that are collected in a Notebook.[56] Be-

[56] Available at https://pedrobrinca.pt/wp-content/uploads/2020/06/BFFHM2020-MCIP.zip.

TABLE A12
REGRESSIONS ON DATA FROM ALESINA ET AL. (2015A), WINSORIZED AT THE 99TH PERCENTILE

Coefficients	(1) Benchmark	(2) Y25/Y75	(3) Y20/Y80	(4) Y10/Y90	(5) Y5/Y95	(6) Y2/Y98	(7) Income Gini
β_1	−0.001	0.006	0.004	−0.002	−0.002	−0.003	0.011
	(0.006)	(0.007)	(0.007)	(0.006)	(0.006)	(0.007)	(0.007)
β_2	−0.003	−0.003	−0.003	−0.001	−0.003	0.000	−0.001
	(0.005)	(0.007)	(0.008)	(0.007)	(0.006)	(0.006)	(0.007)
γ		−2.238**	−1.245	−0.014	0.042	0.021	−1.104***
		(1.018)	(0.764)	(0.343)	(0.135)	(0.050)	(0.386)
ι_1		−1.256*	−0.743	0.119	0.062	−0.012	−0.482**
		(0.644)	(0.531)	(0.232)	(0.078)	(0.030)	(0.210)
ι_2		−0.391	−0.250	−0.107	−0.017	0.022	−0.125
		(0.647)	(0.520)	(0.246)	(0.091)	(0.026)	(0.177)
Constant	0.014***	0.167**	0.117*	0.015	0.003	0.008	0.435***
	(0.005)	(0.070)	(0.064)	(0.050)	(0.034)	(0.014)	(0.148)
Observations	84	84	84	84	84	84	84
R^2	0.006	0.109	0.068	0.011	0.025	0.020	0.160
Number of country1	12	12	12	12	12	12	12

NOTES: The table displays the results from estimating the regression in Equation (3) on data from Alesina et al. (2015a) and measures of income inequality from the EU-SILC. Y25/Y75, Y20/Y80, Y10/Y90, Y5/Y95, and Y2/Y98 represent the share of income of the top 25%, 20%, 10%, 5%, and 2% divided by the share of the bottom 25%, 20%, 10%, 5%, and 2%. Standard errors in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, and $^{*}p < 0.1$.

TABLE A13
BLANCHARD AND LEIGH (2013) REGRESSIONS FOR THE COUNTRIES IN SECTION 7

Coefficients	(1) Blanchard–Leigh	(2) Y25/Y75	(3) Y20/Y80	(4) Y10/Y90	(5) Y5/Y95	(6) Y2/Y98	(7) Income Gini
β	−1.430***	−1.161***	−1.170***	−1.204***	−1.286***	−1.259***	−1.378***
	(0.182)	(0.131)	(0.134)	(0.172)	(0.164)	(0.176)	(0.204)
γ		−0.490	−0.303	−0.033	0.031	0.048	0.365**
		(0.750)	(0.534)	(0.165)	(0.034)	(0.046)	(0.120)
ι		−0.122	−0.119	−0.073	−0.039*	−0.040**	−0.053
		(0.187)	(0.134)	(0.047)	(0.017)	(0.017)	(0.063)
Constant	1.207*	4.304	3.553	1.712	0.616	0.228	−12.831**
	(0.567)	(5.167)	(4.552)	(2.709)	(1.233)	(0.840)	(4.458)
Observations	13	13	13	13	13	13	13
R^2	0.715	0.755	0.750	0.736	0.736	0.748	0.919

NOTES: The table displays the results from estimating the regression in (1) just for the countries in Section 7. These are the countries for which we have enough data to calibrate the model. Y25/Y75, Y20/Y80, Y10/Y90, Y5/Y95, and Y2/Y98 represent the share of income of the top 25%, 20%, 10%, 5%, and 2% divided by the share of the bottom 25%, 20%, 10%, 5%, and 2%. Robust standard errors in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, and $^{*}p < 0.1$.

low we outline the main results of these exercises. We first estimate the ρ without fixed effects, namely we run a pooled OLS:

$$u_{it} = constant + \rho u_{it-1} + \epsilon_{it}. \tag{A.14}$$

In this case, we obtain $\rho = 0.73$, which is closer to other authors, for example, Chang and Kim (2006), who perform a similar exercise on wage rate data from the PSID. Of course in the case of unobserved heterogeneity, the pooled OLS is biased.

Second, we change the sample to keep only individuals with at least 20 observations. This decreases the number of observations substantially, but leads to higher estimates of ρ. Running the fixed effects estimator, we find a $\rho = 0.43$, which, as expected, is higher. Whether this value represents a better estimate is an open question and beyond the scope of our article.

TABLE A14

BLANCHARD AND LEIGH (2013) REGRESSIONS WITH PERCENTAGE OF CREDIT-CONSTRAINED AGENTS AND CONSUMER CREDIT (AS PERCENTAGE OF GROSS INCOME)

VARIABLES	(1) Blanchard–Leigh	(2) Credit-constrained	(3) Consumer Credit
β	−1.095***	−1.181***	−2.239***
	(0.255)	(0.260)	(0.504)
γ		−0.003	−0.005
		(0.006)	(0.017)
ι		0.002	0.003
		(0.004)	(0.010)
Constant	0.775*	0.537	4.885**
	(0.383)	(0.539)	(2.125)
Observations	26	12	13
R^2	0.496	0.800	0.677

NOTES: The table displays the results from estimating the regression in (1) with percentage of credit-constrained agents and consumer credit and interaction of both these variables with the consolidation shocks. The consumer credit is measured as the total household credit excluding mortgages as a percentage of gross income. Robust standard errors in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, and $^{*}p < 0.1$.

TABLE A15

BLANCHARD AND LEIGH (2013) REGRESSIONS WITH HOUSEHOLD DEBT AND INCOME INEQUALITY MEASURES

VARIABLES	(1) Y25/Y75	(2) Y20/Y80	(3) Y10/Y90	(4) Y5/Y95	(5) Y2/Y98	(6) Income Gini
β	−1.682***	−1.654***	−1.405***	−1.275***	−1.389***	−1.960***
	(0.455)	(0.444)	(0.347)	(0.295)	(0.305)	(0.367)
γ_1	−0.020**	−0.019**	−0.017**	−0.014**	−0.016***	−0.016*
	(0.007)	(0.007)	(0.006)	(0.005)	(0.005)	(0.008)
ι_1	0.010**	0.010**	0.006	0.004	0.004	0.007*
γ_2	−1.040*	−0.809*	−0.345	−0.069	0.005	0.179
	(0.509)	(0.412)	(0.210)	(0.063)	(0.056)	(0.254)
ι_2	0.043	0.005	−0.093	−0.084*	−0.081**	0.055
	(0.357)	(0.292)	(0.128)	(0.041)	(0.037)	(0.120)
	(0.004)	(0.004)	(0.004)	(0.003)	(0.003)	(0.004)
Constant	13.148***	12.717***	10.707***	7.452***	5.937***	−0.823
	(3.963)	(3.881)	(3.257)	(1.678)	(1.415)	(10.013)
Observations	25	25	25	25	25	25
R^2	0.672	0.673	0.694	0.699	0.666	0.644

NOTES: The table displays the results from augmenting the regression in Blanchard and Leigh (2013) with household debt γ_1 and an interaction term between household debt and planned fiscal consolidation ι_1 while controlling for income inequality γ_2 and its interaction term with fiscal consolidation ι_2. Robust standard errors in parentheses. Robust standard errors in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, and $^{*}p < 0.1$.

TABLE A16

CONSOLIDATIONS AND CONSUMER CREDIT (AS PERCENTAGE OF GROSS INCOME) FOR THE COUNTRIES IN SECTION 7

VARIABLES	(1) Precrisis household debt
β	−2.239***
	(0.504)
γ	−0.005
	(0.017)
ι	0.003
	(0.010)
Constant	4.885**
	(2.125)
Observations	13
R^2	0.677

$^{***}p < 0.01$, $^{**}p < 0.05$, and $^{*}p < 0.1$. Robust standard errors in parentheses.

TABLE A17

LABOR SUPPLY RESPONSE, WEALTH, AND FISCAL CONSOLIDATION SHOCKS

VARIABLES	(1) $\Delta \ln h$	(2) $\Delta \ln h$	(3) $\Delta \ln h$
β_1	0.137***	0.136***	
	(0.023)	(0.023)	
β_2	−0.042	−0.046	
	(0.046)	(0.046)	
β_3	−0.020*	−0.013	−0.002
	(0.011)	(0.011)	(0.011)
β_4		−0.011	−0.014
		(0.057)	(0.057)
β_5		−0.219***	−0.204***
		(0.069)	(0.070)
Constant	−0.169***	−0.169***	−0.149***
	(0.009)	(0.009)	(0.007)
Observations	72,969	72,969	72,969
R^2	0.121	0.121	0.124
Household FE	Yes	Yes	Yes
Year FE	No	No	Yes

NOTES: The table displays the results from estimating the regression in Equation (16) with total wealth net of debt including home equity. Robust standard errors in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, and $^{*}p < 0.1$.

TABLE A18

CORRELATIONS BETWEEN INCOME INEQUALITY MEASURES AND CREDIT-CONSTRAINED HOUSEHOLDS

VARIABLES	(1) Percentage Constrained
Y2/Y98	−0.2471
Y5/Y95	−0.1161
Y10/Y90	−0.1617
Y20/Y80	−0.2609
Y25/Y75	−0.2800
Income Gini	−0.3550

Our baseline specification is a dynamic panel data with fixed effects. The presence of the lagged dependent variable can introduce a bias in the standard fixed effect estimator (the Nickel's bias) if the time dimension is not sufficiently large. To have a sense of the strength of the bias, we perform a Monte Carlo experiment using the size of our sample $T = 30$ (and some properties of the distributions of our data). Details on the Monte Carlo can be found in the Notebook. We report the average absolute bias and the average variance as a function of the time dimension.

Table A3 shows that while the Nickel's bias is very large for panels with small T, our sample size is sufficiently large to be not too concerned. With $T = 30$, the negative bias is about 12%, implying that a better estimate of our benchmark ρ would be 0.38. In the case of using individuals with a minimum of 20 observations, the implied value, taking into account the bias, would be 0.48.

Admittedly our specification is somewhat different from other works in the literature that focused on the income process estimation. The first difference is that our dependent variable is the hourly wage and not the level of labor income. The second difference is in the specification. To be more precise, a common specification in the literature of the second stage regression is

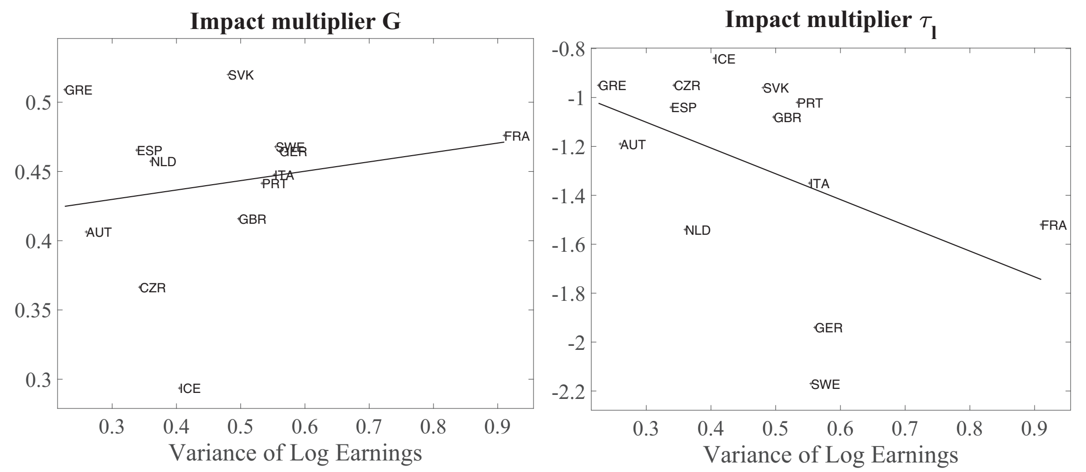
$$\hat{u}_{i,a,t} = \mu_i a + z_{i,a,t} + \zeta_{i,a,t} + \epsilon_{i,a,t},$$

TABLE A19

ACTUAL FISCAL CONSOLIDATION FOR SELECTED COUNTRIES

Country	Actual Consolidation
Austria	1.0
Czech Republic	2.1
France	1.2
Germany	0.3
Greece	10.3
Iceland	4.0
Italy	0.2
Netherlands	0.1
Portugal	2.7
Spain	1.5
Slovakia	2.0
Sweden	0.9
United Kingdom	3.0

NOTES: The table displays actual fiscal consolidations undertaken during the years 2010–2011. Positive values represent a consolidation. All values are in percentage of GDP.



NOTES: The left panel displays the cross-country data for a consolidation implemented through decreasing G (correlation coefficient 0.1868 , $p = 0.5413$), whereas the right panel displays the cross-country data for a consolidation implemented through increasing the labor tax (correlation coefficient -0.4924; $p = 0.0873$).

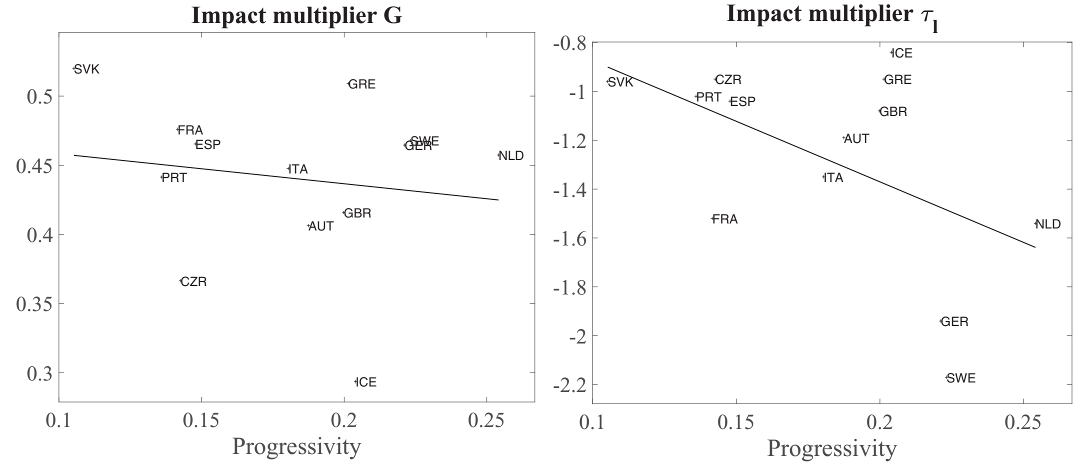
FIGURE A8

IMPACT MULTIPLIER AND VARIANCE OF DISPOSABLE LABOR EARNINGS

where $z_{i,a,t}$ is the transitory component and $\zeta_{i,a,t}$ is the persistent component. Now in our simplified exercise. we put $z = 0$ and think of $\zeta_{i,a,t} = \rho\zeta_{i,a-1,t-1} + \eta^p_{i,a,t}$. The regression we run is $\hat{u}_{i,t} = \alpha_i + \psi\hat{u}_{i,t-1} + v_{i,t}$. The most salient differences are: first, that we do not allow for individual heterogeneity to depend on age (or experience) and second, that the implicit structural (in a time series sense) process for $\zeta_{i,a,t}$ is more complex in our specification (this can be shown with a few steps of algebra). Our choices are motivated by the model specification that does not allow heterogeneity across age or experience but only across a fixed genetic ability.

Finally, an article that uses the same specification as we use is Lillard and Willis (1978). Although in a shorter sample, they find $\rho = [0.35, 0.406]$ depending on the specification.

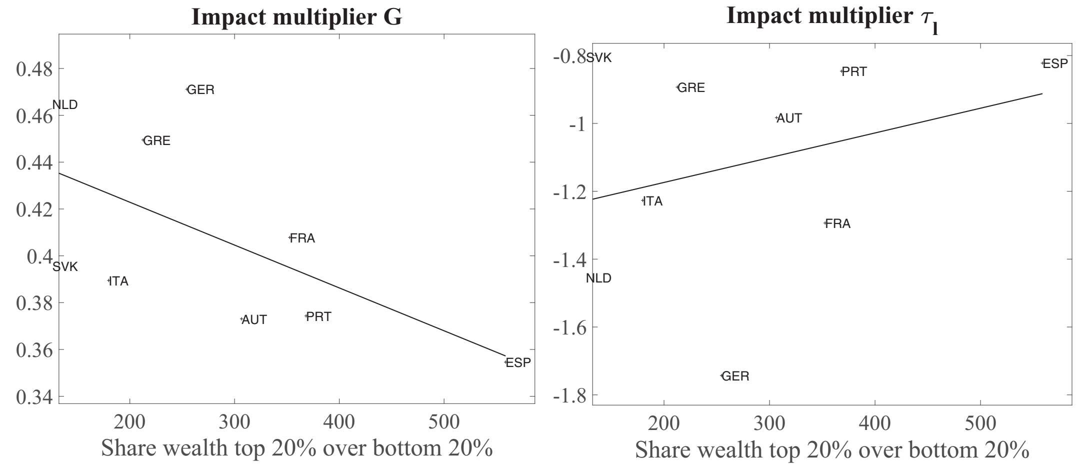
A.10 *The Labor Supply Response to Fiscal Shocks—Additional Evidence.* To study the response of labor supply to fiscal shocks we conduct two additional exercises using the data on fiscal consolidation episodes from Alesina et al. (2015a). First, in Subsection A.10.1 we



NOTES. The left panel displays the cross-country data for a consolidation implemented through decreasing G (correlation coefficient -0.1209, $p = 0.6961$), whereas the right panel right panel displays the cross-country data for a consolidation implemented through increasing the labor tax (correlation coefficient -0.4209; $p = 0.1521$).

FIGURE A9

IMPACT MULTIPLIER AND θ_1



NOTES: On the left panel, we have the cross-country data for a consolidation done by decreasing G (correlation coefficient -0.5833; $p = 0.1080$), whereas on the right panel we have the cross-country data for a consolidation done by increasing the labor tax (correlation coefficient 0.25 00; $p = 0.5206$).

FIGURE A11

IMPACT MULTIPLIER AND WEALTH GINI COEFFICIENT, CALIBRATING EACH COUNTRY TO THE DISTRIBUTION OF LIQUID WEALTH

study the interaction between income inequality and the response of aggregate hours to fiscal consolidation shocks. Then in Subsection A.10.2 we use age as a proxy for wealth to study the labor supply response to fiscal consolidation by household wealth.

A.10.1 *Aggregate hours worked and income inequality.* To investigate how the labor supply response depends on income inequality, we use the Alesina et al. (2015a) data set and hours worked per capita from OECD, from 2007 until 2012. We estimate the following equation:

$$H_{i,t} = \alpha + \beta_1 e^u_{i,t} + \beta_2 e^a_{i,t} + \gamma I_{i,t-1} + \iota_1 e^u_{i,t}(I_{i,t-1} - \mu_I) + \iota_2 e^a_{i,t}(I_{i,t-1} - \mu_I) + \delta_i + \omega_t + \epsilon_{i,t}, \tag{A.15}$$

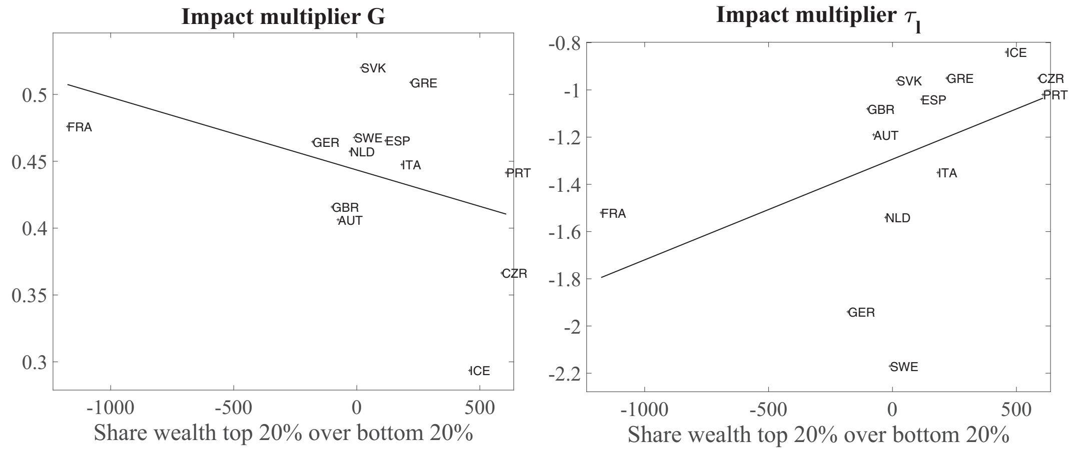


FIGURE A12

IMPACT MULTIPLIER AND WEALTH INEQUALITY MEASURES AS SHARE OF TOP 20% OVER SHARE OF BOTTOM 20%
NOTES: ON THE LEFT PANEL, WE HAVE THE CROSS-COUNTRY DATA FOR A CONSOLIDATION DONE BY DECREASING G (CORRELATION COEFFICIENT 0.2967; $P = 0.3243$), WHEREAS ON THE RIGHT PANEL WE HAVE THE CROSS-COUNTRY DATA FOR A CONSOLIDATION DONE BY INCREASING THE LABOR TAX (CORRELATION COEFFICIENT 0.7143; $P = 0.0081$).

TABLE A20
50% RISK + 50% ABILITY EXPERIMENT VERSUS BENCHMARK

	σ_u		σ_a		Percentage Constrained	
	50–50%	Benchmark	50–50%	Benchmark	50–50%	Benchmark
Austria	0.310	0.180	0.330	0.420	8.7	10.8
Czech Republic	0.300	0.150	0.320	0.420	9.9	7.4
France	0.460	0.510	0.480	0.420	3.4	2.8
Germany	0.420	0.440	0.440	0.420	3.7	3.4
Greece	0.300	0.120	0.310	0.420	4.0	5.8
Iceland	0.350	0.290	0.370	0.420	7.3	6.8
Italy	0.320	0.370	0.340	0.420	2.0	3.9
Netherlands	0.340	0.260	0.360	0.420	3.4	4.9
Portugal	0.390	0.380	0.410	0.420	3.4	5.4
Spain	0.330	0.240	0.350	0.420	4.2	4.5
Slovakia	0.360	0.330	0.390	0.420	2.8	2.8
Sweden	0.400	0.410	0.430	0.420	4.0	3.4
United Kingdom	0.390	0.380	0.410	0.420	3.9	4.0

NOTES: Values for σ_u and σ_a under two different calibration regimes. The "50–50%"—where σ_a and σ_u explain half each of the variance of log wages, net of what is explained by the age profile—and our benchmark calibration. Note that both calibrations yield a very similar pattern of credit-constrained agents (correlation of 0.90; $p < 0.001$).

where $H_{i,t}$ is normalized annual hours worked per capita[57] in country i in year t. The right-hand side of the equation is the same as in Equation (3). The results are presented in Table A4 and establish that labor supply is more responsive to fiscal consolidations in countries with higher inequality, just as our mechanism suggests. Note that, similar to the results in Subsection 3.2, it is the interaction with the unanticipated fiscal consolidations that is statistically significant. For a country with income inequality 1 percentage point above the sample mean, the drop in hours worked is larger by ι_1. Increasing the share of income in the top 10% over the share in the bottom 10% by one standard deviation causes hours worked to drop by 124% more than for a country with average inequality.

[57] We follow Brinca et al. (2016a) and express hours worked as a share of the working day and compute it by using OECD data and multiplying hours worked per employee by total employment, per capita, divided by 5,200.

TABLE A21
COUNTRY-SPECIFIC CALIBRATION TARGETS

	Macro Ratios		Labor Targets		Income Profile Parameters	Taxes			
	K/Y	B/Y	$\bar{n}$	Var($\ln w$)	$\gamma_1, \gamma_2, \gamma_3$	θ_1, θ_2	$\tilde{\tau}_{ss}, \tau_{ss}$	τ_c	τ_k
Austria	3.359	0.432	0.226	0.199	0.155, −0.004, 3.0E-05	0.939, 0.187	0.217, 0.181	0.196	0.240
Czech Republic	6.203	0.206	0.236	0.174	0.174, −0.004, 3.0E-05	0.988, 0.143	0.350, 0.125	0.182	0.220
France	3.392	0.559	0.184	0.478	0.384, −0.008, 6.0E-05	0.915, 0.142	0.434, 0.135	0.183	0.355
Germany	3.013	0.489	0.189	0.354	0.176, −0.003, 2.3E-05	0.881, 0.221	0.206, 0.210	0.155	0.233
Greece	3.262	1.038	0.230	0.220	0.120, −0.002, 1.3E-05	1.062, 0.201	0.280, 0.160	0.154	0.160
Iceland	4.334	0.213	0.308	0.249	0.161, −0.003, 1.9E-05	0.868, 0.204	0.055, 0.000	0.253	0.200
Italy	3.943	0.893	0.200	0.225	0.114, −0.002, 1.4E-05	0.897, 0.180	0.329, 0.092	0.145	0.340
Netherlands	2.830	0.232	0.200	0.282	0.307, −0.007, 4.9E-05	0.938, 0.254	0.102, 0.200	0.194	0.293
Portugal	3.229	0.557	0.249	0.298	0.172, −0.004, 2.6E-05	0.937, 0.136	0.238, 0.110	0.208	0.234
Spain	3.378	0.368	0.183	0.225	0.114, −0.002, 1.4E-05	0.904, 0.148	0.305, 0.064	0.144	0.296
Slovakia	3.799	0.317	0.204	0.250	0.096, −0.002, 1.7E-05	0.974, 0.105	0.326, 0.131	0.181	0.151
Sweden	2.155	−0.034	0.233	0.315	−0.021, 0.001, −1.2E-05	0.796, 0.223	0.326, 0.070	0.255	0.409
United Kingdom	2.315	0.371	0.231	0.302	0.183, −0.004, 2.2E-05	0.920, 0.200	0.105, 0.090	0.163	0.456

NOTES: Macro ratios: K/Y is derived from Penn World Table 8.0, average from 1990-2011; B/Y is the average of net public debt from 2001 to 2008 (IMF). Labor targets: $\bar{n}$ is hours worked per capita derived from OECD data, average from 1990–2011; Var($\ln w$) and $\gamma_1, \gamma_2, \gamma_3$ are from the most recent LIS survey available before 2008 using the gross wage of primary employment series. Italy only reports net wages, which we then modify with our tax progressivity function. Data from Portugal come from Quadros de Pessoal 2009 database. Taxes: θ_1, θ_2 are as discussed in Subsection A.2; $\tilde{\tau}_{ss}, \tau_{ss}$ are the average social security withholdings faced by the average earner (OECD) from 2001 to 2007; τ_k and τ_c are either taken from Trabandt and Uhlig (2011) or calculated using their approach, representing average effective tax rates from 1995 to 2007. τ_k for Iceland comes from the Iceland Ministry of Industries and Innovation.

TABLE A22
COUNTRY-SPECIFIC PARAMATER VALUES ESTIMATED BY SMM

Country	β_1	β_2	β_3	χ	b	σ_u	φ
Austria	0.959	1.003	0.964	14.40	0.00	0.176	4.30
Czech Republic	0.999	1.041	0.996	21.00	0.00	0.145	11.70
France	0.957	1.013	0.990	18.03	0.25	0.506	3.24
Germany	0.952	0.997	0.952	16.93	0.09	0.439	3.60
Greece	0.989	0.997	0.969	16.50	0.00	0.121	3.70
Iceland	0.962	0.996	0.962	7.53	0.08	0.294	9.60
Italy	0.992	1.016	0.984	20.30	0.00	0.237	6.00
Netherlands	0.942	0.986	0.973	14.75	0.15	0.263	2.98
Portugal	0.960	0.991	0.960	11.50	0.00	0.380	5.20
Spain	0.970	0.997	0.983	24.47	0.00	0.237	5.00
Slovakia	0.984	0.993	0.984	20.40	0.00	0.326	7.20
Sweden	0.917	0.971	0.944	9.40	0.33	0.407	2.20
United Kingdom	0.939	0.968	0.939	12.40	0.10	0.379	4.90

TABLE A23
PARAMETERS HELD CONSTANT ACROSS COUNTRIES

Parameter Preferences	Value	Description	Source
η	1	Inverse Frisch Elasticity	Trabandt and Uhlig (2011)
σ	1.2	Risk aversion parameter	Literature
Technology			
α	0.33	Capital share of output	Literature
δ	0.06	Capital depreciation rate	Literature
ρ	0.335	$u' = \rho u + \epsilon, \quad \epsilon \sim N(0, \sigma_\epsilon^2)$	PSID 1968–1997
σ_a	0.423	Variance of ability	European economies average from Brinca et al. (2016b)

TABLE A24

BLANCHARD AND LEIGH (2013) REGRESSIONS WITH MODEL GENERATED FORECAST ERRORS FOR τ_l-CONSOLIDATIONS

Coefficients	(1) Blanchard–Leigh	(2) Y25/Y75	(3) Y20/Y80	(4) Y10/Y90	(5) Y5/Y95	(6) Y2/Y98	(7) Income Gini
τ_l consolidation	−2.578	−2.321***	−2.662***	−4.134***	−5.987***	−5.603***	−2.475***
	(2.143)	(0.656)	(0.657)	(0.691)	(0.825)	(1.134)	(0.656)
Inequality		−2.144***	−1.500***	−0.549**	−0.189	−0.143	−0.880***
		(0.524)	(0.377)	(0.187)	(0.117)	(0.089)	(0.187)
Interaction		−0.006	−0.219	−0.563***	−0.635***	−0.407***	0.148
		(0.341)	(0.243)	(0.117)	(0.099)	(0.092)	(0.117)
Constant	−4.263**	5.694*	4.208	1.389	−0.422	−0.554	22.246***
	(1.508)	(2.621)	(2.443)	(2.173)	(2.065)	(1.989)	(5.478)
Observations	13	13	13	13	13	13	13
R^2	0.129	0.657	0.649	0.646	0.634	0.599	0.784

NOTES: The table displays the results from running the regression in Equation (2), using model generated forecast errors for tax-based consolidations. Y25/Y75, Y20/Y80, Y10/Y90, Y5/Y95, and Y2/Y98 represent the share of income of the top 25%, 20%, 10%, 5%, and 2% divided by the share of the bottom 25%, 20%, 10%, 5%, and 2%. Robust standard errors in parentheses. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

As mentioned, the mechanism relies on the labor supply of credit-constrained agents falling much less than for unconstrained agents. This is consistent with the empirical evidence on labor supply in Domeij and Floden (2006), and Subsection 8.2.

A.10.2 *Age as proxy for wealth.* To provide evidence on the relationship between wealth and labor supply responses to fiscal consolidation shocks in European countries, we conduct an experiment analogous to Subsection 3.2 using household level data. Instead of explaining output responses, we study the labor supply response of employed workers to fiscal consolidation shocks, and instead of income inequality we look at the impact of age (which can be thought of as a proxy for wealth). The data are obtained from the EU LFS household data, which does not have wealth. The dependent variable is *usual* weekly hours worked by worker i in country c and year t. We look at the impact of fiscal events that Alesina et al. (2015a) identify as exogenous interacted with worker age (measured by membership in five year cohorts, which is one of the limiting conditions of the anonymous data). The data are a repeated cross-section. We use roughly the same time frame as the experiment in Section 3.2, that is, 2008–2014, and estimate the following model:

$$\begin{aligned} h_{i,c,t} &= \alpha + \beta_1 e^u_{i,c,t} + \beta_2 e^a_{i,c,t} + \gamma age_{i,c,t} + \iota_1 e^u_{i,c,t}(age_{i,c,t}) + \iota_2 e^a_{i,c,t}(age_{i,c,t}) \\ &\quad + \delta_c + \omega_t + \epsilon_{i,c,t}, \end{aligned} \tag{A.16}$$

where e^u and e^a are the fiscal adjustment episodes identified in Alesina et al. (2015a), (unanticipated and anticipated). First, running the unconditional regression (i.e., without worker age and the interactions), we get a negative impact of the size of the consolidation on hours worked. However, including the age cohort interactions shows that there is heterogeneity in response by age. The regression shows that young workers do not reduce their hours like older workers (and in fact the youngest cohorts even increase their hours).

This is consistent with our mechanism, since younger (as a proxy for poorer) workers show a lower labor supply elasticity with respect to the consolidation shock than older (richer) workers (see Figure A5). Furthermore this is consistent with the empirical evidence provided in Section 8: poorer agents are more inelastic with respect to the consolidation shock than richer agents. The fact that we obtain similar results with different data sets and different identification strategies is reassuring.

A.11 *Additional Figures and Tables.*

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