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How Important Are Credit Constraints?

by

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Government Size and Business Cycle Volatility; How Important Are Credit Constraints?

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Abstract

In this paper we analyze how the availability of credit influences the relationship between government size as a proxy for fiscal stabilization policy and the amplitude of business cycle fluctuations in a sample of advanced OECD countries. Interpreting relatively low loan-to-value ratios as an indication for tight credit constraints, we find that government size exerts a stabilizing effect on output and consumption growth fluctuations only when credit constraints are relatively tight. Our results are robust with respect to different measures of government size and provide support for the hypothesis that credit market frictions play a crucial role in the transmission of fiscal policy.

Keywords: Business cycle, volatility, fiscal policy, stabilization policy

JEL codes: E62, E32

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1 Introduction

Can fiscal policy contribute to macroeconomic stability? This question has a long tradition in the theoretical as well as empirical literature and has received renewed attention in the aftermath of the 2007-2009 recession.[1] Galí (1994) and Fatás and Mihov (2001) were among the first to empirically show that countries characterized by high ratios of government spending to GDP tend to have less volatile business cycles. Since government size is found to be positively correlated with the extent to which automatic stabilizers operate (see e.g. Dolls et al., 2012; Girouard and André, 2005; Van den Noord, 2000), these empirical results suggest that fiscal policy indeed exerts a stabilizing effect on the business cycle, at least if it is conducted through automatic stabilizers.

Theoretically, however, the effect of automatic stabilizers on the business cycle is less clear. Although there is little doubt that automatic stabilizers, such as income tax and social expenditures, offset fluctuations in disposable incomes, their overall effectiveness in terms of the stabilization of economic activity depends crucially on the response of private demand to fiscal policy actions, which is a more controversial issue. A number of studies argue that the reaction of private demand is closely related to the extent to which credit constraints are binding (Auerbach and Feenberg, 2000; Dolls et al., 2012). Standard models with forward-looking agents and frictionless financial markets predict that private consumption remains unchanged despite changes in taxes and transfers as long as the present value of lifetime disposable income does not change. If, in contrast, credit constraints restrict private consumption, then an increase in current disposable income resulting from, e.g., a tax reduction leads to higher consumption. Thus, fiscal policy should be able to stabilize fluctuations in economic activity via the tax and transfer system much along the lines of traditional Keynesian arguments if the availability of credit is limited.

Fiscal policy may also mitigate fluctuations in disposable incomes through discretionary changes in the tax and transfers system if these changes are implemented in a way that systematically reacts to the business cycle. In addition, discretionary fiscal policy also involves adjustments in government purchases, such as government consumption and investment, which may also dampen business cycle fluctuations. Yet, the effect of government purchases on private consumption also depends on the availability of credit. In models without financial frictions, an increase in government purchases reduces private consumption because of negative wealth effects

[1] See Ramey (2011) and Cwik and Wieland (2011) for recent surveys.

(Linnemann and Schabert, 2003; Baxter and King, 1993) and intertemporal substitution effects (Davig and Leeper, 2011; Woodford, 2010; Christiano et al., 2009; Benassy, 2007). Hence, the ability of fiscal policy to dampen the business cycle via variations in spending is limited in these models. In fact, a fiscal expansion during a recession may even amplify the downturn if wealth and substitution effects are sufficiently strong. Nevertheless, a negative correlation between government size and the volatility of output can still be obtained in these models. However, as Andres et al. (2008) show, such a negative correlation is the consequence of a composition effect, since private consumption and investment actually become more volatile. Thus, in models without financial frictions, a relatively large public sector may coincide with low business cycle volatility simply because public spending itself is not as volatile as private sector demand. To generate a positive response of private consumption to an increase in government spending, Andres et al. (2008) and Galí et al. (2007) include so-called rule-of-thumb agents in addition to forward-looking, optimizing agents in their models. Since rule-of-thumb agents are assumed to neither borrow nor save, they behave in a more Keynesian way in the sense that consumption spending is closely related to current income.[2] This type of rule–of–thumb behavior can be interpreted as the consequence of binding credit constraints or, more generally, limited asset market participation.[3]

To sum up, fiscal policy should be able to dampen business cycle fluctuations, via the stabilization of private demand when credit constraints are binding. Against this background, we empirically explore the relationship between government size, business cycle volatility and credit market imperfections based on a panel of 18 OECD countries from 1970 to 2007. Specifically, we study if and how the influence of government size on the amplitude of fluctuations in output growth depends on the availability of credit. We use the loan-to-value (LTV) ratio, which is the highest mortgage loan that households can get as a fraction of the value of a house. As emphasized by Jappelli and Pagano (1994), LTV ratios provide a measure of financial constraints on households that is comparable across countries (see also Perotti, 1999).

Taking potential endogeneity into account, we find that government size significantly reduces the magnitude of fluctuations in output and consumption growth rates when LTV ratios are

[2]It must be noted however, that the presence of rule-of-thumb agents by itself is not necessarily sufficient to generate an expansionary consumption response of aggregate consumption. While rule-of-thumb behavior reduces the impact of the negative wealth effect, labor income must increase to obtain a positive consumption response. Therefore, as pointed out by Galí et al. (2007), prices have to be sufficiently sticky. Otherwise, the lower marginal labor productivity associated with higher employment leads to a decline in real wage.

[3]Although credit market frictions are perhaps the most prominent interpretation, rule-of-thumb behavior can be motivated in a number of ways, such as buffer-stock savings behavior (Mankiw, 2000) or deviations from rationality in the form of myopia or debt aversion (Thaler, 1992).

low, that is, when credit is relatively tight. When LTV ratios are high, in contrast, government size exerts a positive, albeit insignificant effect. Thus, while we partly confirm the findings in Galí (1994) and Fatás and Mihov (2001), we contribute to the literature by showing that the stabilizing effect of government size is closely related to the availability of credit. This result also provides additional empirical support for the literature that emphasizes the role of financial market frictions for the transmission of fiscal policy.

Our paper is closely related to the branch of the literature that studies the influence of credit market frictions on the transmission of fiscal policy. On the basis of a stochastic general equilibrium (DSGE) model estimated with U.S. data, Bilbiie et al. (2008) argue that increased asset market participation over time has reduced the influence of fiscal policy shocks in the U.S. To analyze the transmission of fiscal policy in the euro area, Forni et al. (2009) estimate a DSGE model featuring rule-of-thumb agents. Auerbach and Feenberg (2000) and Dolls et al. (2012) analyze the effects of automatic stabilizers using a micro-simulation model and conclude that their effectiveness depends strongly on the presence of credit constraints. Perotti (1999) also takes LTV ratios into account when analyzing the effects of fiscal policy on consumption growth. While he is primarily interested in demonstrating that fiscal contractions can have expansionary effects on private consumption in times of fiscal distress, we are interested in the influence of fiscal policy on the amplitude of fluctuations in general. Auerbach and Gorodnichenko (2010) show that fiscal multipliers are larger in recessions than in boom periods. This result is consistent with our findings since credit constraints are more likely to be binding in recessions as argued in Tagkalakis (2008).

The remainder of the paper is structured as follows: in Section 2, we discuss estimation strategy and describe the data set. Section 3 presents our estimation results. Section 4 concludes the paper.

2 Estimation Strategy and Data

Our analysis is based on variants of the following regression:

$$Fluctuation_{it} = \alpha G_{it} + \beta Glob_{it} + \lambda_i + \lambda_t + \epsilon_{it}, \tag{1}$$

where $Fluctuation_{it}$ is a measure of the amplitude of business cycle fluctuations, G_{it} is a proxy for government size, $Glob_{it}$ is a control variable that captures the degree of openness, and λ_i and λ_t are country and year fixed effects, respectively.

We follow Morgan et al. (2004) and construct a measure of the amplitude of fluctuations in real GDP growth based on the estimated residual, $\hat{u}_{it}$, of the regression

$$\Delta \log y_{it} = \nu_i + \nu_t + u_{it}, \tag{2}$$

where y_{it} is real GDP and ν_i and ν_t denote country and year fixed effects respectively. We define the dependent variable in equation (1) as $Fluctuation_{it} = |\hat{u}_{it}|$, which is the size of the deviation of real GDP growth from average growth for a given country-year (see also Kalemli-Ozcan et al., 2010; Thesmar and Thoenig, 2011). Since $Fluctuation_{it}$ varies across countries and also across time, we are able to exploit the panel structure of the data. Thus, here we deviate from Fatás and Mihov (2001) who use the standard deviation of real output growth to measure the size of business cycle fluctuations and limit their analysis to a cross-section of countries. We also estimate variants of equation (1) where we replace the amplitude of fluctuations in real output growth with the amplitude of fluctuations of real consumption growth to determine whether fiscal policy exerts a stabilizing influence on private demand. For these estimations, we construct a measure of the amplitude of real consumption growth fluctuations analogously to output growth fluctuations. Bootstrapped standard errors are reported throughout the paper to account for the construction of $Fluctuation_{it}$.

We measure government size either by the log of the ratio of government spending to GDP, denoted by Gov_{it}, or by the log of tax revenues to GDP, Tax_{it}. While Gov_{it} is frequently used as an indicator of the extent of stabilization policy (see e.g. Fatás and Mihov, 2001), we use Tax_{it} as an additional proxy since government revenues are rather sensitive with respect to the business cycle (see e.g. Auerbach and Feenberg, 2000; Cottarelli and Fedelino, 2010). Although we interpret government size as an indicator for the stabilizing role of fiscal policy, countries characterized by large government sectors may also be exposed to destabilizing fiscal shocks to a greater extent. Fatás and Mihov (2003) show that discretionary policy implemented in a way that is unrelated to macroeconomic conditions increases the volatility of real GDP growth. Nevertheless, as long as fiscal shocks are quantitatively small, the effect of systematic fiscal policy should prevail. Forni et al. (2009) conclude that fiscal policy shocks contribute little to the cyclical variability of macroeconomic variables in the euro area.

We include the log of the KOF index of economic globalization (Dreher, 2006), denoted by $Glob_{it}$, to control for openness. Rodrik (1998) finds that more open countries experience more volatile fluctuations. Using firm-level data, di Giovanni and Levchenko (2009) also conclude that trade openness increases volatility. In contrast, Haddad et al. (2010) argue that openness may also reduce volatility if countries are sufficiently diversified. In addition, Ilzetzki et al. (2010) find that fiscal multipliers are smaller in open economies. By controlling for openness, we also take into account that the effectiveness of fiscal policy may depend on the degree of openness. The KOF index provides a summary measure of the economic dimension of globalization. Note, however, that the KOF index may be endogenous in equation (1) since it captures, among other things, actual economic flows such as foreign direct investment, that may depend on business cycle volatility. To cope with this issue we re-estimate our specifications using only the economic restrictions part of the index. Since these restrictions refer to the institutional and legal environment, they are plausibly exogenous for our purposes. Since the estimation results, which are available upon request, are rather similar to those obtained with the overall index, we rely on $Glob_{it}$ in our main analysis as it captures economic globalization in a broader way.[4]

Our data set comprises 18 OECD countries, listed in Table 1, and covers the period from 1970 to 2007. Real GDP growth rates are taken from the OECD Country Statistical Profiles 2010 database and real private final consumption expenditures from the OECD Economic Outlook database. For Germany, we use consumption data provided by the German Federal Statistical Office (Destatis) for the period before 1991. Government spending series are taken from the OECD Economic Outlook database, where we use data from Andres et al. (2008) to substitute missing values. Tax revenue series come from the OECD Revenue Statistics database. Figure 1 shows that spending and tax revenues as percentages of GDP, averaged over countries, increased over time and the increase is more pronounced for spending than for revenues. Moreover, the increase in spending reversed in the early 1990s because of consolidation measures taken in many European countries.

Note that our sample includes the well documented decline in macroeconomic volatility during the mid 1980s associated with the Great Moderation (see e.g. Stock and Watson, 2005). Since we include time fixed effects, we control for changes in the amplitude of fluctuations that are common to all countries in the sample (see also Coric, 2011, for a discussion of the global dimension of the Great Moderation). Furthermore, since we also include country fixed effects in equation (1), we capture any influence of institutional variables, such as characteristics of the electoral and the political system, which are emphasized in Carmignani et al. (2011).

Government size, measured either by Gov_{it} or Tax_{it}, can to be endogenous in equation (1) since large fluctuations in output growth are likely to trigger fiscal policy responses that result

[4]Potential endogeneity problems are also the reason for why we do not include other control variables which are closely related to GDP as in Fatás and Mihov (2001).

in variations in the ratios of government spending and tax revenues to GDP. To allow for a causal interpretation, we identify the exogenous variation in government size using instrumental variables that are related to structural aspects and are therefore plausibly exogenous with respect to the amplitude of the business cycle. Specifically, we use the log of the urban population as a percentage of the total population, $Urban_{it}$, and the fraction of left-wing parties in parliament, $Left_{it}$ to instrument Gov_{it} and Tax_{it}. While the public finance literature suggests that urbanization is likely to influence the size of governments, the sign of the effect is ambiguous *a priori*. Although countries with larger urban populations may be able to provide public services at a lower cost by exploiting economies of scale (see e.g. Fatás and Mihov, 2001), it is also conceivable that a highly concentrated population leads to congestion in the consumption of public services. Hence, government action to prevent congestion externalities becomes increasingly necessary and, as a consequence, may result in a higher public spending (Buchanan, 1970). For $Left_{it}$, the party ideology hypothesis (see e.g. Le Maux et al., 2010) suggests a positive sign in the first-stage regression since left-wing governments typically spend more than right-wing governments. $Left_{it}$ is defined as the share of votes that socialist, left-socialist and communist parties obtained in the last parliament election. We calculate $Urban_{it}$ based on data provided by the United Nations World Urbanization Prospects database and data for the construction of $Left_{it}$ are taken from Armingeon et al. (2010).[5] Note that our panel is slightly unbalanced because of missing values of $Left_{it}$ for Greece, Portugal and Spain in the early 1970s.

We measure the availability of credit using the LTV ratios reported in Almeida et al. (2006) for the 1970s, 1980s, and 1990s. Since our macroeconomic series run until 2007, we extend the series until the end of our sample with the LTV ratios reported for the 1990s.[6] For Austria, Greece, Portugal, and for Japan for the 1970s, we use data reported in Tagkalakis (2008). As in Jappelli and Pagano (1994) and Perotti (1999) we distinguish between loose and tight credit constraints in the following way: we define a dummy L_{it} as $L_{it} = 1$ if the LTV ratio in country i in year t is at least 80 percent and $L_{it} = 0$ otherwise. Country-years for which $L_{it} = 0$ are considered to be characterized by tight constraints on the availability of credit and country-years with $L_{it} = 1$ are considered to be observations for which constraints are less binding. What we are primarily interested in is the influence of the availability of credit on the relationship between

[5] Except for $Left_it$, all right-hand side variables enter in logs. $Left_it$ enters in levels since some observations are equal to zero.

[6] In a closely related paper, Dolls et al. (2012) proxy credit constraints using variables such as financial wealth, home ownership, and survey outcomes. The availability of these variables is substantially more limited than for the LTV ratio which renders them unsuitable for our analysis.

government size and the size of business cycle fluctuations. To investigate this issue, we estimate equation (1) separately for observations characterized by loose or tight credit constraints. That is, we compare the effect of government size across the two subsamples characterized by either $L_{it} = 0$ or $L_{it} = 1$.

Note that a sample selection problem could arise if $Fluctuation_{it}$ influences the assignment of observations to one of the two groups for which we estimate equation (1). However, since the construction of L_{it} relies on the long-run behavior of the LTV ratios, it is more likely to mirror structural characteristics of the financial system and therefore L_{it} is credibly exogenous with respect to $Fluctuation_{it}$. In fact, Table 1 shows that the assignment of observations into groups of tightly and loosely credit constrained observations is quite stable over time. Although some countries switch between groups, these switches do not appear to be driven by the macroeconomic conditions prevalent at the time of the switch. For instance, several countries switch to the group characterized by relatively loose constraints in the early 1980s, a time of high macroeconomic volatility. It is hard to imagine that banks eased access to credit because of a highly volatile macroeconomic environment.

It still appears conceivable that the degree to which credit constraints bind depends on the average size of fluctuations. Suppose that countries that experience more volatile business cycles on average also tend to be characterized by lower LTV ratios, as lenders adjust their behavior over time. Then countries with relatively pronounced fluctuations in macroeconomic activity would be included in the $L_{it} = 0$ group. In addition, a selection bias could also arise if the construction of L_{it} is driven by variables that are related to both: the size of fluctuations and LTV ratios. In either case, we should observe systematic differences in the size of fluctuations across the two groups. However, in our sample the average magnitude of output growth fluctuations is fairly similar in both groups. The mean of $Fluctuation_{it}$ is 1.247 percentage points for country-years characterized by loose constraints and 1.249 percentage points for country-years with tight constraints.[7] Moreover, a two-sample Kolmogorov-Smirnov test for equality of distributions does not reject the null hypothesis that the realizations of $Fluctuation_{it}$ in both groups of observations are drawn from the same distribution.[8]

While selection problems seem unlikely, we nevertheless test for the presence of a sample selection bias combining the procedures proposed by Lee (1978) and Semykina and Wooldridge (2010): we first estimate a pooled probit regression with L_{it} as the dependent variable (see also

[7] The average annual growth rate of real GDP is slightly below 3 percent in the full sample.

[8] The null hypothesis that the observations in the two subsamples are drawn form the same distribution is not rejected with a p-value of 0.608.

Wooldridge, 2010, p. 833). As explanatory variables we use the exogenous regressor in equation (1), $Glob_{it}$; the excluded instruments $Left_{it}$ and $Urban_{it}$; as well as their country-means. In addition, we also include a dummy variable indicating the legal tradition of country i which we denote by $Civil_i$, to improve the explanatory power of the probit regression. This dummy is defined as $Civil_i = 1$ if country i has a civil law tradition and $Civil_i = 0$ in case of a common law tradition. La Porta et al. (1997) argue that countries with a common law tradition offer systematically better investor protection which fosters the development of financial markets. To the extent that developed financial markets also provide easier access to credit, we expect civil law countries to have lower LTV ratios.[9] Data on the legal tradition are taken from La Porta et al. (1997).

The next step in the testing procedure is to calculate the inverse Mills ratios for country-year pairs with loose and tight constraints: For country-year pairs with loose credit constraints, let $Mills1_{it} = \phi(z'\pi)/\Phi(z'\pi)$, where z is the vector of regressors and π is the vector of estimated parameters of the probit model. $\phi(z'\pi)$ and $\Phi(z'\pi)$ are the density and cumulative probability distribution functions of the standard normal distribution evaluated at $z'\pi$. Similarly, $Mills0_{it} = -\phi(z'\pi)/(1 - \Phi(z'\pi))$ is the inverse Mills ratio for country-years with tight constraints. Finally, we reestimate equation (1) by fixed effects two-stage least squares and include $Mills0_{it}$ or $Mills1_{it}$ as additional regressors. If either of the two inverse Mills ratios turns out to have explanatory power in the second-stage regressions, then the original estimations may suffer from a selection bias.

Given the relatively long time dimension of our panel relative to the cross-sectional dimension, non-stationarity of the series could be of concern. We determine the integration properties of the variables using the panel unit root test of the Phillips-Perron-Fisher type (see e.g. Breitung and Pesaran, 2005). This test is appropriate for our dataset since it relies on T asymptotics with fixed N and allows for unbalanced panels. To account for a limited amount of cross-sectional dependence, we subtract the cross-sectional mean of each variable. Since the $Fluctuation_{it}$ variables do not trend over time, we only include country fixed effects in the regressions. For the remaining variables, we include a time trend and country fixed effects. We test the null hypothesis of a unit root using different lag lengths, that is, with different orders of residual autocorrelation. We set the maximum lag length to 4, which roughly corresponds to $T^{1/3}$ (see Said and Dickey, 1984). Table 2 shows that we can reject the null hypothesis of a unit root at least at the 10-percent level in all but one case. For Gov_{it}, the significance level is 14.3 percent

[9]Almeida et al. (2006) also relate financial development to LTV ratios.

when only one residual lag is considered. For higher lag lengths, the test also rejects the null hypothesis also for Gov_{it} at standard levels of significance. Overall, these results indicate that the series are stationary.

3 Estimation Results

Table 3 presents the results for basic specification (1) using government spending as a percentage of GDP, Gov_{it}, to measure government size. Column (I) shows the results for the full sample and Columns (II) and (III) display the results for country-year pairs with LTV ratios of at least 80 percent (Column II) or below 80 percent (Column III).

From Column (I) we see that Gov_{it} exerts the expected dampening influence on output growth fluctuations in the full sample, which is in line with the results reported in Fatás and Mihov (2001). While Fatás and Mihov (2001) use a different measure of volatility and estimate a cross-section regression, they report estimated coefficients which are of a similar order of magnitude. The control variable $Glob_{it}$ is positively signed, but insignificant. The first-stage results are rather satisfactory. The instruments $Left_{it}$ and $Urban_{it}$ are both highly significant and enter the first-stage with the expected signs. $Left_{it}$ exerts a positive effect on Gov_{it}, which is in line with the party ideology hypothesis and the positive effect of $Urban_{it}$ is consistent with the idea that the provision of public services is more expensive in urban areas because of congestion. The Hansen J-test does not reject the null hypothesis that the instruments are uncorrelated with the error term in the second-stage regression, suggesting that our instruments are valid. Since we obtain a (bootstrapped) F-statistic for the excluded instruments of 61.25, we also consider our instruments to be strong. Note also that $Glob_{it}$ exerts a significantly positive effect in the first-stage regression, which supports Rodrik (1998) who argues that more open economies have larger governments.

We are mainly interested in how the influence of government size differs across observations characterized by loosely or tightly binding credit constraints, that is, high and low LTV ratios. Comparing Columns (II) and (III) shows that the dampening effect of Gov_{it} is present only in the subsample comprising country-years characterized by tight constraints. In contrast, when credit constraints are loose, fiscal policy has a positive, but insignificant influence on the magnitude of output fluctuations.

These effects of government size on output growth volatility are quantitatively substantial. Suppose the share of government expenditures in GDP increases by 10 percent. Such an increase would raise the average share of government spending in GDP in the $L_{it} = 0$ subsample from 44.6 percent to 49 percent. Since the estimated coefficients in Table 3 are semi-elasticities this increase in Gov_{it} reduces $Fluctuation_{it}$ by 0.37 percentage points, which corresponds to roughly 30 percent of the average amplitude of output growth fluctuations in the subsample characterized by binding credit constraints. In contrast, if credit constraints are relatively loose, a ten-percent increase in the share of government expenditures in GDP increases average output volatility by about 17.5 percent.[10]

The first-stage regression results are similar to those reported in Column (I). Although $Left_{it}$ is insignificant in Column (II) and $Urban_{it}$ is insignificant in Column (III), both instrumental variables remain positively signed, and the F-test and J-test statistics indicate that the instruments are strong and valid in both subsamples.

Does our estimation suffer from a sample selection bias? To analyze this issue, we test for a selection bias using the procedure described in Section 2. Column (I) in Table 4 shows the estimation results for the probit regression with L_{it} as the dependent variable. $Glob_{it}$ is highly significant in this estimation, albeit negatively signed, which is somewhat surprising as one would expect that highly globalized countries provide better access to credit. $Urban_{it}$ and $Left_{it}$ are both negatively signed, but only $Urban_{it}$ is significant. Finally, $Civil_i = 1$ significantly reduces the probability that the LTV ratio in country i is at least 80 percent. Thus, countries with a civil law tradition have a significantly higher probability of being characterized by tighter constraints. Assuming that access to credit is more restricted in countries with less developed financial systems, this result is consistent with La Porta et al. (1998) who argue that countries with civil law legal traditions tend to have less developed financial systems. Columns (II) and (III) show that the inverse Mills ratios obtained from the probit estimation are insignificant in the second stages in both subsamples, indicating the absence of a sample selection bias.

As additional robustness checks, for which detailed estimation results are available upon request, we also estimate the basic specification of dropping single years and single countries and find that our conclusions do not change. Similarly, weighting countries by their populations does not change the results.

Table 5 presents the results for Tax_{it} as an alternative measure of government size. Columns (I) to (III) show that using Tax_{it} does not change our conclusions with respect to credit constraints. Recall from Figure 1, that government spending and tax revenues as percentages of

[10] According to Table 3, a one-percent increase in government size reduces $Fluctuation_{it}$ by 0.0373 percentage points if $L_{it} = 0$, and increases $Fluctuation_{it}$ by 0.0219 percentage points if $L_{it} = 1$. The sample means of $Fluctuation_{it}$ are 1.25 for the $L_{it} = 0$ and $L_{it} = 1$ observations, respectively.

GDP evolve somewhat differently over time. Nevertheless, we obtain rather similar results with both proxies for government size. Yet, quantitatively, the effects of government size are now more pronounced regardless of the tightness of credit constraints. This result is not unexpected, since tax revenues are highly responsive to the business cycle and the ratio of tax revenues to GDP may therefore capture the response of fiscal policy to the business cycle to a greater extent than Gov_{it}. Although the F-test statistic is below the rule-of-thumb value of 10 suggested by Staiger and Stock (1997), in the subsample for $L_{it} = 1$, the first-stage regression results are in line with the results presented in Column (II) in Table 3.[11]

The last years included in our sample, from the early 2000s until the onset of the global financial crisis during the summer of 2007, were characterized by an abundance of liquidity and rather loose financial conditions on a global level. As it seems plausible that these factors increased the availability of credit to an extent that may not be fully captured by the LTV ratios, our results might be influenced by these exceptional financial conditions. To see if this is indeed the case, we exclude these years from the estimation sample. The estimation results for the shorter sample in Table 6, for Gov_{it} in Columns (I) to (III) and for Tax_{it} in Columns (IV) to (VI) confirm our main finding that the stabilizing effect of fiscal policy depends on the availability of credit. We also see that the dampening effect of government size proxied by either Gov_{it} or Tax_{it} obtained for $L_{it} = 0$ observations is larger when we exclude the most recent years. This outcome is consistent with our assertion that exceptionally loose credit constraints in recent years reduced the stabilizing effect of government size.

While the 2000s were exceptional with respect to financial conditions, the 1990s were characterized by important institutional changes in a number of the countries included in our sample. For EU member countries, the Treaty of Maastricht stipulates the Excessive Deficit Procedure (EDP), which established numerical fiscal deficit and debt rules as a prerequisite for membership in the European Monetary Union (EMU). In light of the EDP several EU countries introduced or renewed fiscal rules during the 1990s (Debrun et al., 2008).[12] Changes in the institutional framework within which fiscal policy operates were not limited to EU member countries. Japan, for instance, adopted a new fiscal rule in 1996, and in 1997 a new fiscal spending act was passed to reduce public deficits and expenditure growth (Von Hagen, 2006).

[11] We also tested for a selection bias in this specification with Tax_{it}. Results are overall similar to those found with Gov_{it} and do not indicate the presence of a selection bias.

[12] For instance, real expenditure ceilings and rules with respect to the allocation of excess revenues were introduced in the Netherlands in 1994. In Austria, a National Stability Pact was stipulated in 1999 to ensure compliance with the EDP and the Treaty of Amsterdam.

One could argue that these institutional changes may have changed the relationship between fiscal policy and the business cycle and might therefore influence our results. Although the ability of fiscal policy to counteract business cycle fluctuations may have become increasingly limited because of the introduction of fiscal rules, leading to more volatile cycles, fiscal rules may also have decreased the size of fluctuations by reducing the magnitude of discretionary policy shocks.[13] To eliminate any potential influences that the institutional changes implemented during the 1990s may have on our analysis, we shorten the sample period further to end in 1991, the year before the EDP was stipulated.

Table 7 shows that the estimation results for this sample also support our main conclusion that government size exerts a dampening effect only when credit constraints are tight (Columns (III) and (VI)). Although the coefficients are significant only at the 13- and 11-percent significance levels in the $L = 0$ subsample, they are in a similar order of magnitude as in the corresponding columns of Table 6. Thus, it appears that the institutional changes implemented during the 1990s had only a limited influence on the relationship between government size and output volatility. The exceptionally loose financial conditions that prevailed during the 2000s had a relatively larger influence.

Finally, it is still possible that our estimations pick up a composition effect. Do countries with larger government sectors experience smaller fluctuations simply because the public sector is less volatile than the private sector? An alternative interpretation is that fiscal policy manages to dampen fluctuations in economic activity by exerting a stabilizing influence on private sector demand. Note that the composition effect should operate independently of LTV ratios. In this sense, our results presented thus far already suggest that we do not pick up a composition effect because the effect of Gov_{it} turns out to be closely related to the LTV ratio. Nevertheless, to provide additional evidence, we reestimate equation (1) with the volatility of real consumption growth as the dependent variable and either Gov_{it} or Tax_{it} as a proxy for governrnment size. If fiscal policy exerts a stabilizing influence via private demand, then we should also observe a negative relationship between government size and the volatility of real consumption growth rates in countries with tight credit constraints.

We see from Table 8 that government size, measured by either Gov_{it} (Columns (I) to (III)), or by Tax_{it} (Columns (IV) to (VI)), exerts a dampening effect on consumption growth fluctuations only for country-years with relatively tight credit constraints. The low first stage F-statistic in

[13] Fatás and Mihov (2006) show for federated states in the U.S. that the second effect dominates and that fiscal rules have dampened state business cycles.

Column (V) signals that the estimation with Tax_{it} likely suffers from weak instruments when the sample characterized by loose constraints is considered. Overall, however, these results support the interpretation that in cases of tight credit constraints, fiscal policy manages to stabilize private sector demand, which, in turn, feeds back to economic activity and results in smoother business cycles.

4 Summary and Concluding Remarks

In this paper, we study how the availability of credit influences the stabilizing influence of government size on the business cycle. We essentially combine two strands of the existing literature: the first studies the influence of government size on the volatility of fluctuations in economic activity and the second stresses credit market frictions as a crucial element for the transmission of fiscal policy. We find that credit market frictions indeed play a key role. While government size exerts a statistically and economically significant dampening effect on output growth fluctuations when credit is tight, government size may even be associated with more pronounced business cycles when credit is readily available. These results are fully consistent with the theoretical prediction that credit market frictions, which make demand strongly dependent on current income, are essential for fiscal policy to exert a stabilizing influence.

Based on estimates of the fiscal multiplier, Ilzetzki et al. (2010) conclude that the effectiveness of fiscal policy has declined over time owing to increasing trade integration and a less accommodating monetary policy stance. Our results provide a complementary reason for the decline in the effectiveness of fiscal policy over time, namely increased asset market participation and a readier access to credit. In our sample, six of the 12 countries characterized by tight credit constraints in the 1970s show increasing LTV ratios over time (see Table 1). Only one country (Sweden) shows a decline in its LTV ratio. Hence, given our results, this trend toward greater credit availability may be another reason why fiscal multipliers have declined over time.

Finally, although we find that larger governments exert a dampening effect on output volatility, it should be kept in mind that the overall welfare implications of larger governments are not easy to evaluate. While smoother business cycles should be welfare improving, recent analysis (see e.g. Folster and Henrekson, 2001; Uhlig, 2010) documents that pairing larger governments with unfavorable expenditure and tax structures, may have adverse consequences on the long-run growth performance of an economy.

References

Almeida, H., Campello, M., Liu, C., 2006. The financial accelerator: Evidence from international housing markets. Review of Finance 10 (3), 321–352.

Andres, J., Domenech, R., Fatás, A., 2008. The stabilizing role of government size. Journal of Economic Dynamics and Control 32 (2), 571–593.

Armingeon, K., Engler, S., Potolidis, P., Gerber, M., Leimgruber, P., 2010. Comparative political data set 1960-2008. Institute of Political Science, University of Berne.

Auerbach, A. J., Feenberg, D., 2000. The significance of federal taxes as automatic stabilizers. Journal of Economic Perspectives 14 (3), 37–56.

Auerbach, A. J., Gorodnichenko, Y., 2010. Measuring the output responses to fiscal policy. NBER Working Papers 16311, National Bureau of Economic Research, Inc.

Baxter, M., King, R. G., 1993. Fiscal policy in general equilibrium. American Economic Review 83 (3), 315–34.

Benassy, J.-P., 2007. IS-LM and the multiplier: A dynamic general equilibrium model. Economics Letters 96 (2), 189–195.

Bilbiie, F. O., Meier, A., Müller, G. J., 2008. What accounts for the changes in U.S. fiscal policy transmission? Journal of Money, Credit and Banking 40 (7), 1439–1470.

Breitung, J., Pesaran, M., 2005. Unit roots and cointegration in panels. Cambridge Working Papers in Economics 0535, Faculty of Economics, University of Cambridge.

Buchanan, J. M., 1970. The public finances, 3rd Edition. Irwin, Homewood, Illinois.

Carmignani, F., Colombo, E., Tirelli, P., 2011. Macroeconomic risk and the (de)stabilising role of government size. European Journal of Political Economy 27 (4), 781–790.

Christiano, L., Eichenbaum, M., Rebelo, S., 2009. When is the government spending multiplier large? NBER Working Papers 15394, National Bureau of Economic Research, Inc.

Coric, B., 2011. The global extent of the Great Moderation. Oxford Bulletin of Economics and Statistics forthcoming.

Cottarelli, C., Fedelino, A., 2010. Automatic stabilizers and the size of government: Correcting a common misunderstanding. IMF Working Papers 10/155, International Monetary Fund.

Cwik, T., Wieland, V., 2011. Keynesian government spending multipliers and spillovers in the euro area. Economic Policy 26 (67), 493–549.

Davig, T., Leeper, E. M., 2011. Monetary-fiscal policy interactions and fiscal stimulus. European Economic Review forthcoming.

Debrun, X., Moulin, L., Turrini, A., Ayuso-i-Casals, J., Kumar, M. S., 2008. Tied to the mast? National fiscal rules in the European Union. Economic Policy 23, 297–362.

di Giovanni, J., Levchenko, A. A., 2009. Trade openness and volatility. The Review of Economics and Statistics 91 (3), 558–585.

Dolls, M., Fuest, C., Peichl, A., 2012. Automatic stabilizers and economic crisis: US vs. Europe. Journal of Public Economics 96 (3-4), 279–294.

Dreher, A., 2006. Does globalization affect growth? Evidence from a new index of globalization. Applied Economics 38 (10), 1091–1110.

Fatás, A., Mihov, I., 2001. Government size and automatic stabilizers: International and intranational evidence. Journal of International Economics 55 (1), 3–28.

Fatás, A., Mihov, I., 2003. The case for restricting fiscal policy discretion. The Quarterly Journal of Economics 118 (4), 1419–1447.

Fatás, A., Mihov, I., 2006. The macroeconomic effects of fiscal rules in the US states. Journal of Public Economics 90 (1-2), 101–117.

Folster, S., Henrekson, M., 2001. Growth effects of government expenditure and taxation in rich countries. European Economic Review 45 (8), 1501–1520.

Forni, L., Monteforte, L., Sessa, L., 2009. The general equilibrium effects of fiscal policy: Estimates for the euro area. Journal of Public Economics 93 (3-4), 559–585.

Galí, J., 1994. Government size and macroeconomic stability. European Economic Review 38 (1), 117–132.

Galí, J., López-Salido, J. D., Vallés, J., 2007. Understanding the effects of government spending on consumption. Journal of the European Economic Association 5 (1), 227–270.

Girouard, N., André, C., 2005. Measuring cyclically-adjusted budget balances for oecd countries. OECD Economics Department Working Papers 434, OECD Publishing.

Haddad, M. E., Lim, J. J., Saborowski, C., 2010. Trade openness reduces growth volatility when countries are well diversified. Policy Research Working Paper Series 5222, The World Bank.

Ilzetzki, E., Mendoza, E. G., Végh, C. A., 2010. How big (small?) are fiscal multipliers? NBER Working Papers 16479.

Jappelli, T., Pagano, M., 1994. Saving, growth, and liquidity constraints. The Quarterly Journal of Economics 109 (1), pp. 83–109.

Kalemli-Ozcan, S., Sorensen, B. E., Volosovych, V., 2010. Deep financial integration and volatility. University Economic Research Forum Working Papers 1006, TUSIAD-Koc University Economic Research Forum.

La Porta, R., Lopez-de Silanes, Shleifer, A., Vishny, R. W., 1998. Law and finance. Journal of Political Economy 106, 1113–55.

La Porta, R., Lopez-de Silanes, F., Shleifer, A., Vishny, R. W., 1997. Legal determinants of external finance. Journal of Finance 52 (3), 1131–50.

Le Maux, B., Rocaboy, Y., Goodspeed, T., 2010. Political fragmentation, party ideology and public expenditures. Public Choice forthcoming.

Lee, L.-F., 1978. Unionism and wage rates: A simultaneous equations model with qualitative and limited dependent variables. International Economic Review 19 (2), 415–33.

Linnemann, L., Schabert, A., 2003. Fiscal policy in the new neoclassical synthesis. Journal of Money, Credit and Banking 35 (6), 911–29.

Mankiw, N. G., 2000. The savers-spenders theory of fiscal policy. American Economic Review 90 (2), 120–125.

Morgan, D., Rime, B., Strahan, P. E., 2004. Bank integration and state business cycles. The Quarterly Journal of Economics 119 (4), 1555–1584.

Perotti, R., 1999. Fiscal policy in good times and bad. The Quarterly Journal of Economics 114 (4), 1399–1436.

Ramey, V. A., 2011. Can government purchases stimulate the economy? Journal of Economic Literature 49 (3), 673–85.

Rodrik, D., 1998. Why do more open economies have bigger governments? Journal of Political Economy 106 (5), 997–1032.

Said, E., Dickey, D. A., 1984. Testing for unit roots in autoregressive-moving average models of unknown order. Biometrika 71 (3), 599–607.

Semykina, A., Wooldridge, J. M., 2010. Estimating panel data models in the presence of endogeneity and selection. Journal of Econometrics 157 (2), 375–380.

Staiger, D., Stock, J. H., 1997. Instrumental variables regression with weak instruments. Econometrica 65 (3), 557–586.

Stock, J. H., Watson, M. W., 2005. Understanding dhanges in international business cycle dynamics. Journal of the European Economic Association 3 (5), 968–1006.

Tagkalakis, A., 2008. The effects of fiscal policy on consumption in recessions and expansions. Journal of Public Economics 92 (5-6), 1486–1508.

Thaler, R., 1992. The winners curse: Anomalies and paradoxes of economic life. Free press, New York.

Thesmar, D., Thoenig, M., 2011. Contrasting trends in firm volatility. American Economic Journal: Macroeconomics 3 (4), 143–80.

Uhlig, H., 2010. Some fiscal calculus. American Economic Review 100 (2), 30–34.

Van den Noord, P., 2000. The size and role of automatic fiscal stabilizers in the 1990s and beyond. OECD Economics Department Working Papers 230, OECD Publishing.

Von Hagen, J., 2006. Fiscal rules and fiscal performance in the European Union and Japan. Monetary and Economic Studies 24 (1), 25–60.

Woodford, M., 2010. Simple analytics of the government expenditure multiplier. NBER Working Papers 15714.

Wooldridge, J. M., 2010. Econometric analysis of cross section and panel data. Vol. 2 of MIT Press Books. The MIT Press.

Table 1: Loose and Tight Credit Constraints

	Loose constraints: $L_{it} = 1$	Tight constraints: $L_{it} = 0$
Australia	1980-2007	1970-1979
Austria		1970-2007
Belgium	1990-2007	1970-1989
Canada	1980-2007	1970-1979
Finland	1970-2007	
France	1970-2007	
Germany	1990-2007	1970-1989
Greece		1970-2007
Ireland	1970-2007	
Italy		1970-2007
Japan		1970-2007
Netherlands		1970-2007
Norway	1980-2007	1970-1979
Portugal		1970-2007
Spain	1980-2007	1970-1979
Sweden	1970-1989	1990-2007
United Kingdom	1970-2007	
United States	1970-2007	

Notes: Loan-to-value ratios that exceed 80 percent indicate loose credit constraints. The grouping of observations is based on LTV ratios reported in Almeida et al. (2006). For Austria, Greece, and Portugal and for Japan for the 1970s , we use the LTV ratios reported in Tagkalakis (2008).

Table 2: Fisher-Phillips-Perron-type Panel Unit Root Test (p-values)

	Number of Included Lags			
	1	2	3	4
$Fluctuation_{it}$ (Output growth)	0.0000	0.0000	0.0000	0.0000
$Fluctuation_{it}$ (Consumption growth)	0.0000	0.0000	0.0000	0.0000
Gov_{it}	0.1429	0.0960	0.0709	0.0681
Tax_{it}	0.0278	0.0269	0.0274	0.0327
$Glob_{it}$	0.0583	0.0588	0.0534	0.0646
$Left_{it}$	0.0282	0.0153	0.0119	0.0265
$Urban_{it}$	0.0000	0.0000	0.0000	0.0000

Notes: The table shows the p-values for the inverse χ^2 test statistic with 38 degrees of freedom. The top line gives the number of lags included. The null hypothesis is that the series contain unit roots. For the $Fluctuation_{it}$ variables, country fixed effects are included in the regressions. For all other variables, a time trend and country fixed effects are included.

Table 3: Government Size, Credit Constraints and Output Growth Fluctuations

	(I) Full Sample	(II) Loose Constraints $L_{it} = 1$	(III) Tight Constraints $L_{it} = 0$
Gov_{it}	-2.7392*	2.1939	-3.7300*
	(1.5377)	(2.4738)	(2.1059)
$Glob_{it}$	1.2367	-2.1005	1.3412
	(1.0829)	(2.1942)	(1.1774)
Obs	668	358	310
First stage regression results			
$Glob_{it}$	0.5875***	0.7142***	0.3756***
	(0.0546)	(0.0721)	(0.1130)
$Left_{it}$	0.0028***	0.0007	0.0048***
	(0.0008)	(0.0009)	(0.0011)
$Urban_{it}$	0.7455***	1.0656***	0.1026
	(0.1115)	(0.1707)	(0.2156)
F-statistic	61.25	19.72	34.52
adjusted R^2	0.8190	0.7879	0.8734
OID (p-value)	0.2860	0.2148	0.4647

Notes: The table shows 2SLS estimation results for the full sample in Column (I), for country-years characterized by loose constraints, that is, LTV ratios that exceed 80 percent in Column (II), and for country-years characterized by tight constraints, that is LTV ratios below 80 percent in Column (III). The dependent variable is the magnitude of output growth fluctuations, $Fluctuation_{it}$. Government size, Gov_{it} is instrumented with the urban population as a percentage of the total population, $Urban_{it}$, and the fraction of left-wing parties in parliament, $Left_{it}$. Bootstrapped standard errors in brackets. ***, **, and * denote statistical significance at the 1-percent, 5-percent, and 10-percent level. The table reports the F-statistic of the excluded instruments and adjusted R^2 for the first-stage estimation. OID (p-value) is the p-value associated with the Hansen J-test of the over-identifying restrictions. All specifications include country and year fixed effects.

Table 4: Testing for Sample Selection Bias

	(I)	(II)	(III)
	Probit	*Second Stage Regression Results*	
		Loose constraints	Tight constraints
		$L_{it} = 1$	$L_{it} = 0$
Gov_{it}		2.6706	-3.5598
		(3.2293)	(3.3057)
$Glob_{it}$	-4.1618***	-2.0784	0.2627
	(0.8911)	(2.0945)	(1.9991)
$Urban_{it}$	-6.5785***		
	(2.0784)		
$Left_{it}$	-0.0059		
	(0.0108)		
$Civil_i$	-1.9346***		
	(0.2230)		
$Mills1_{it}$		-0.2362	
		(0.6952)	
$Mills0_{it}$			0.5973
			(1.3822)
OID (p-value)		0.2149	0.3276
Observations	668	358	310
(Pseudo) R^2	0.3122		
		First Stage Regression Results	
$Glob_{it}$		0.6716***	-1.2913***
		(0.0894)	(0.2341)
$Urban_{it}$		1.0064***	-1.3829***
		(0.2222)	(0.2398)
$Left_{it}$		0.0007	0.0007
		(0.0009)	(0.0009)
$Mills1_{it}$		0.0277	
		(0.0476)	
$Mills0_{it}$			-0.7557***
			(0.0988)
F-statistic		10.39	19.41
adjusted R^2		0.7875	0.8999

Notes: Notes: Column (I) shows probit estimation results. The dependent variable is equal to one for country-years when the LTV ratio exceed 80 percent, and equal to zero otherwise. Columns (II) and (III) show 2SLS estimation results for country-years characterized by loose constraints, that is LTV ratios of at least 80 percent in Column (II), and for country-years characterized by tight constraints, that is LTV ratios below 80 percent in Column (III). The dependent variable is the magnitude of output growth fluctuations, $Fluctuation_{it}$. Government size, Gov_{it} is instrumented with the urban population as a percentage of the total population, $Urban_{it}$, and the fraction of left-wing parties in parliament, $Left_{it}$. $Civil_i$ is a dummy variable that is equal to one if country i has civil law legal tradition. Bootstrapped standard errors are in brackets. ***, **, and * denote statistical significance at the 1-percent, 5-percent, and 10-percent level. The table reports the F-statistic of the excluded instruments and adjusted R^2 for the first-stage estimation. OID (p-value) is the p-value associated with the Hansen J-test of the over-identifying restrictions. $Mills1_{it}$ and $Mills0_{it}$ are the inverse Mills ratios from the Probit regression summarized in Column (I). All specifications include country and year fixed effects.

Table 5: Government Size Measured as Tax Revenues to GDP

	(I) Full Sample	(II) Loose Constraints $L_{it} = 1$	(III) Tight Constraints $L_{it} = 0$
Tax_{it}	-4.2769	4.2459	-5.1480*
	(2.7773)	(6.7806)	(3.1002)
$Glob_{it}$	1.6512	-2.1294	3.2108*
	(1.4571)	(3.0690)	(1.9219)
Obs	668	358	310
First stage regression results			
$Glob_{it}$	0.4677***	0.3855***	0.6273***
	(0.0450)	(0.0516)	(0.0919)
$Left_{it}$	0.0016**	0.0006	0.0032***
	(0.0007)	(0.0007)	(0.0010)
$Urban_{it}$	0.4816***	0.4413***	0.1355
	(0.1073)	(0.1338)	(0.1967)
F-statistic	18.76	5.83	13.94
adjusted R^2	0.8965	0.9168	0.9051
OID (p-value)	0.2365	0.2027	0.3952

Notes: The table shows 2SLS estimation results for full sample, for country-years characterized by loose constraints, that is, LTV ratios that exceed 80 percent, and for country-years characterized by tight constraints, that is LTV ratios below 80 percent. The dependent variable is the magnitude of output growth fluctuations. Government size, Tax_{it}, is instrumented with the urban population as a percentage of the total population, $Urban_{it}$, and the fraction of left-wing parties in parliament, $Left_{it}$. Bootstrapped standard errors are in brackets. ***, **, and * denote statistical significance at the 1-percent, 5-percent, and 10-percent level. The table reports the F-statistic of the excluded instruments and adjusted R^2 for the first-stage estimation. OID (p-value) is the p-value associated with the Hansen J-test of the over-identifying restrictions. All specifications include country and year fixed effects.

Table 6: Government Size, Credit Constraints and Output Growth Fluctuations, Sample 1970-2000

	(I) Full Sample	(II) Loose Constraints $L_{it} = 1$	(III) Tight Constraints $L_{it} = 0$	(IV) Full Sample	(V) Loose Constraints $L_{it} = 1$	(VI) Tight Constraints $L_{it} = 0$
Gov_{it}	-3.9096*	3.4982	-7.1070**			
	(2.2742)	(2.3001)	(3.4436)			
Tax_{it}				-6.2873**	5.2373	-6.9575**
				(3.0839)	(4.8184)	(3.1227)
$Glob_{it}$	0.9948	-4.2540*	1.8522	1.5330	-3.3722	4.6539**
	(1.3987)	(2.3091)	(1.3667)	(1.5411)	(2.3042)	(1.9306)
Obs	542	281	261	542	281	261
First stage regression results						
$Glob_{it}$	0.4760***	0.7033***	0.1283	0.3764***	0.3231***	0.5378***
	(0.0743)	(0.0855)	(0.1338)	(0.0514)	(0.0561)	(0.0998)
$Left_{it}$	0.0023**	0.0002	0.0042***	0.0024***	0.0010	0.0045***
	(0.0010)	(0.0010)	(0.0013)	(0.0007)	(0.0007)	(0.0011)
$Urban_{it}$	0.8994***	1.4682***	0.1105	0.5651***	0.7605***	0.0210
	(0.1441)	(0.1848)	(0.2460)	(0.1197)	(0.1365)	(0.2252)
F-statistic	34.95	32.80	16.22	24.35	15.55	20.80
adjusted R^2	0.8361	0.8051	0.8963	0.9130	0.9231	0.9278
OID (p-value)	0.1154	0.2064	0.7269	0.1600	0.1420	0.8589

Notes: The table shows 2SLS estimation results for full sample, for country-years characterized by loose constraints, that is, LTV ratios that exceed 80 percent, and for country-years characterized by tight constraints, that is LTV ratios below 80 percent. The dependent variable is the magnitude of output growth fluctuations. Government size, Gov_{it} and Tax_{it}, is instrumented with the urban population as a percentage of the total population, $Urban_{it}$, and the fraction of left-wing parties in parliament, $Left_{it}$. Bootstrapped standard errors are in brackets. ***, **, and * denote statistical significance at the 1-percent, 5-percent, and 10-percent level. The table reports the F-statistic of the excluded instruments and adjusted R^2 for the first-stage estimation. OID (p-value) is the p-value associated with the Hansen J-test of the over-identifying restrictions. All specifications include country and year fixed effects.

Table 7: Government Size, Credit Constraints and Output Growth Fluctuations, Sample 1970-1991

	(I) Full Sample	(II) Loose Constraints $L_{it}=1$	(III) Tight Constraints $L_{it}=0$	(IV) Full Sample	(V) Loose constraints $L_{it}=1$	(VI) Tight Constraints $L_{it}=0$
Gov_{it}	-4.2588	1.0686	-6.5965			
	(3.0668)	(3.0560)	(4.3451)			
Tax_{it}				-5.7631	0.8867	-7.7667
				(4.0345)	(5.0364)	(4.8048)
$Glob_{it}$	-1.6006	-1.8083	-0.9707	-0.4652	-1.6545	2.8092
	(1.2436)	(1.7420)	(2.1025)	(1.5709)	(1.7657)	(2.8092)
Obs	380	182	298	380	182	198
First stage regression results						
$Glob_{it}$	0.0153	0.3198***	-0.3562***	0.2077***	0.1695**	0.1898**
	(0.0965)	(0.1028)	(0.1311)	(0.0636)	(0.0774)	(0.0958)
$Left_{it}$	0.0036***	0.0047***	0.0021**	0.0026***	0.0034***	0.0019**
	(0.0010)	(0.0009)	(0.0011)	(0.0007)	(0.0008)	(0.0008)
$Urban_{it}$	1.0481***	1.4911***	0.9851	0.7752***	0.7616***	0.8070***
	(0.1884)	(0.1966)	(0.2599)	(0.1488)	(0.1706)	(0.2000)
F-statistic	25.55	39.91	12.57	23.24	15.44	16.55
adjusted R^2	0.9006	0.8629	0.9361	0.9429	0.9281	0.9583
OID (p-value)	0.8566	0.3906	0.7381	0.8428	0.4166	0.6910

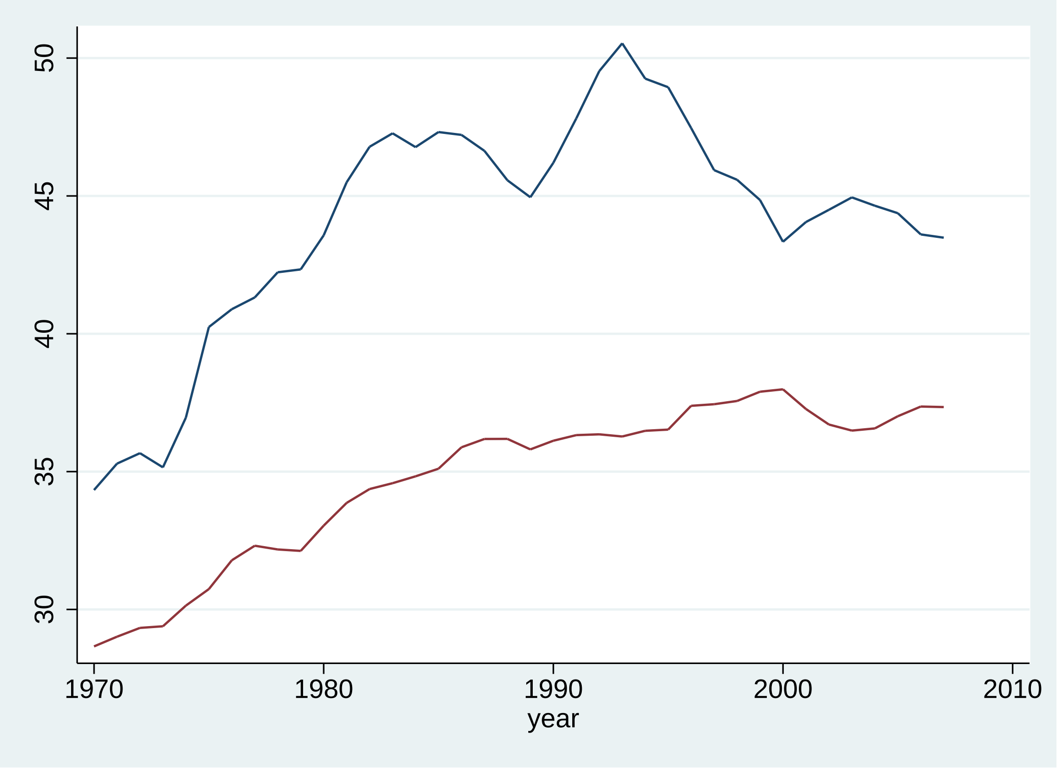
Notes: The table shows 2SLS estimation results for full sample, for country-years characterized by loose constraints, that is LTV ratios of at least 80 percent, and for country-years characterized by tight constraints, that is LTV ratios below 80 percent. The dependent variable is the magnitude of output growth fluctuations. Government size, Gov_{it} and Tax_{it}, is instrumented with the urban population as a percentage of the total population, $Urban_{it}$, and the fraction of left-wing parties in parliament, $Left_{it}$. Bootstrapped standard errors in brackets. ***, **, and * denote statistical significance at the 1-percent, 5-percent, and 10-percent level. The table reports the F-statistic of the excluded instruments and adjusted R^2 for the first-stage estimation. OID (p-value) is the p-value associated with the Hansen J-test of the over-identifying restrictions. All specifications include country and year fixed effects.

Table 8: Consumption Growth Fluctuations

	(I) Full Sample	(II) Loose Constraints $L_{it}=1$	(III) Tight Constraints $L_{it}=0$	(IV) Full Sample	(V) Loose Constraints $L_{it}=1$	(VI) Tight Constraints $L_{it}=0$
Gov_{it}	-2.8072**	2.9124	-3.6327**			
	(1.2521)	(3.0474)	(1.6281)			
Tax_{it}				-4.4997**	9.0640	-5.2017**
				(2.1926)	(12.0067)	(2.2846)
$Glob_{it}$	1.5951*	-2.1392	1.5433	2.1172*	-3.6013	3.5356**
	(0.9117)	(2.5531)	(0.9854)	(1.2029)	(5.1696)	(1.5291)
OID (p-value)	0.4741	0.4357	0.4971	0.4178	0.4904	0.6472
Observations	653	352	301	653	352	301
First Stage Regression Results						
$Glob_{it}$	0.5815***	0.7032***	0.3583***	0.4743***	0.3890***	0.6350***
	(0.0548)	(0.0701)	(0.1085)	(0.1093)	(0.0497)	(0.0878)
$Left_{it}$	0.0029***	0.0003	0.0049***	0.0017***	0.0002	0.0033***
	(0.0008)	(0.0009)	(0.0011)	(0.0007)	(0.0007)	(0.0010)
$Urban_{it}$	0.7461***	0.9988***	0.0711	0.4743***	0.3097***	0.1363
	(0.1104)	(0.1792)	(0.2270)	(0.1093)	(0.1172)	(0.1967)
F-statistic	61.37	15.98	31.73	18.19	3.57	14.99
adjusted R^2	0.8144	0.7881	0.8696	0.9033	0.7881	0.9041

Notes: The table shows 2SLS estimation results for full sample, for country-years characterized by loose constraints, that is LTV ratios of at least 80 percent, and for country-years characterized by tight constraints, that is, LTV ratios that exceed 80 percent. The dependent variable is the magnitude of consumption growth fluctuations. Government size is measured by either Gov_{it} or Tax_{it} which are instrumented with the urban population as a percentage of the total population, $Urban_{it}$, and the fraction of left-wing parties in parliament, $Left_{it}$. Bootstrapped standard errors are in brackets. ***, **, and * denote statistical significance at the 1-percent, 5-percent, and 10-percent level. The table reports the F-statistic of the excluded instruments and adjusted R^2 for the first-stage estimation. OID (p-value) is the p-value associated with the Hansen J-test of the over-identifying restrictions. All specifications include country and year fixed effects.

Figure 1: Average Government Size



Notes: The solid line is average government spending as a percentage of GDP and the broken line is average tax revenues as a percentage of GDP. Each series is constructed as the cross-section average at each point in time for the period from 1970 to 2007.

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