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FULL ARTICLE

Government spending and credit market: Evidence from Italian (NUTS 3) provinces

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Abstract

This study examines the effects of government spending shocks on the Italian credit market using NUTS 3 data over the sample period 2011–2018. The empirical methodology is based on a local projection IV and the identification of a public spending shock is achieved by constructing a Bartik instrument. The empirical evidence shows a mild positive effect of 1% increase in government spending relative to GDP on the growth of the volume loans relative to GDP. However, the empirical findings show that government spending does not help to ameliorate neither the "size bias," that is the financial constraints which small firms face relative to larger ones, nor the "home bias" in lending related to the process of bank consolidation in Italy.

KEYWORDS
Bartik instrument, credit, government spending, local projection

JEL CLASSIFICATION
E62, G21

1 | INTRODUCTION

In a scenario of low interest rates, depicting the current economic conditions, the role of fiscal policy in revitalizing credit markets in economic downturns (beyond unconventional monetary policies) has gained attention from scholars and policymakers.

The focus of this study is on the fiscal policy effects (in terms of government spending shocks) on credit market. As pointed by Auerbach et al. (2020) there is no consensus on the effects of government spending shock on credit markets.

Traditional Keynesian theory suggests absence of an increase in interest rates related to an increase of government spending in a liquidity trap regime, given that the associated rise in the money demand is fully matched by liquidity abundance. Neo-Keynesian models, emphasize the role played by the expected inflation channel. For instance, Christiano et al. (2011) show that, in a liquidity trap regime, fiscal policy shock reduces real interest rates, through an increase in expected inflation. Murphy and Walsh (2022) rationalize a zero or negative impact on interest rates associated with a government spending increase by showing that the latter implies an increase in bond demand (due to a rise in aggregate income) exceeding the government needs to borrow to pay for the spending. Auerbach et al. (2020), points to two transmission channels. First, an increase in local government spending can be interpreted as a wealth transfer, given that it refers to the component of outlays derived from prior contract obligations which can be anticipated. This transmission channel works especially in segmented loan markets, lowering the cost of credit for the local economy given an improvement in the balance sheet of private sector borrowers. The second channel is interpreted as new production. This component is not anticipated, and it is associated with a reduction in the likelihood of a local recession, thereby, implying a further reduction of banks' risk profile assessment of local borrowers and a further reduction in the local cost of credit. This second transmission channel works even in case of integrated local markets. As pointed by Auerbach et al. (2020), this mechanism is akin to the financial accelerator emphasized in Bernanke et al. (1999) and also by post-Keynesian macroeconomic theory that introduces the idea of money supply endogeneity (Dow, 1996; Palley, 2002, among the others) determined through loans provided by banks which in turn generate new deposits.

As pointed out by Chick and Dow (1988), adopting Kaldor's theory, the supply of money is endogenous not only at the national level but also at the regional one, and this is relevant for our analysis that uses regional data. Moreover, post-Keynesian macroeconomic theory emphasizes liquidity preference of banks and of investors providing funds to them as additional channels affecting credit rationing. Government spending can, therefore, diminish not only the probability of an economy to switch into recession, but also the liquidity preference of banks, associated to their need to adjust capital requirements, given the lower risk weights associated with bank assets and lower devaluation of collateral backing loans. Furthermore, government spending can also exert a downward pressure on the liquidity preference of funds providers for banks by reducing the perceived bank credit portfolio risk. This effect can reduce the cost of funding for banks. Finally, an increase in government spending could reduce the liquidity preference of banks (hence credit rationing) especially when loans have to be allocated to categories of borrowers for which constraints, triggered by information asymmetries, are binding especially during economic downturns. These categories are small firms and also borrowers located in a geographical area where a large share of loans is provided by banks whose headquarters are distantly located.[1]

The contribution of this study to the literature on the empirical analysis of fiscal policy impact on credit markets is twofold. First, while the focus of the previous studies is on the US, we concentrate on the (average) impact of government spending on the local Italian credit markets. Second, we explore whether government spending can have an impact on two features related to the Italian credit market. More specifically, we study whether government spending shocks can ameliorate, first, a bias related to the firm size (small firms face more credit constraints than the remaining ones). Small firms generate a considerable share of overall value added in the Italian non-financial business economy. Second, we assess whether government spending can ameliorate the home bias related to the credit constraints affecting borrowers located distantly from the headquarters of the bank supplying credit. This is motivated by the banking consolidation process that has taken place over the last three decades in Italy, which has given banks headquartered in northern Italy a central position in the national credit market.

The empirical analysis is based on a local projection instrumental variables (IV) method. More specifically, we employ a two-stage estimation method applied to NUTS 3 data. In the first stage, the identification of the exogenous

[1]Small firms credit rationing triggered by liquidity preference of banks is highlighted by Dow (1996). More generally, Palley (2002, 2017) links liquidity preference of banks to their management of assets (and liabilities), occurring through a reallocation of borrowers across credit risk categories. A locally targeted government spending increase can, therefore, imply a credit portfolio rebalancing towards borrowers located in the region object of the policy intervention.

variation in government spending is achieved by constructing a Bartik (1991) type instrument, which according to Auerbach et al. (2020), allows us to retrieve the unanticipated, new production component of government spending, which is the main driver of the relationship between public spending and bank assessment of borrowers' risk profiles. In the second stage we estimate a panel regression to obtain local projection (see Jordà, 2005) of credit to the identified government spending shock. The analysis of the impact of government spending on size bias is based on the response of credit to three different categories of borrowers: Non-Financial Corporations and Producer Households to represent the aggregate of businesses, firms with less than 20 employees and producer households to represent the small businesses, and consumer households to represent families. The analysis of the impact of government spending on home bias relies on splitting the sample according to the geographical location of borrowers and creditors, focusing on two main macro-regions: Centre–North and South of Italy ("Mezzogiorno").

The rest of the paper is organized as follows. Section 2 provides a literature review on the empirical studies of the impact of government spending on credit markets. Section 3 describes the data, the empirical methodology, including the identification of government spending, and the empirical evidence. Section 4 describes the robustness analysis and Section 5 gives some concluding remarks.

2 | LITERATURE REVIEW

Standard Keynesian and neoclassical theories argue that an increase in government spending leads to a contraction of the credit market as it causes interest rates to rise. Therefore, "Government spending has traditionally been considered a counterproductive tool for stimulating credit" (Auerbach et al., 2020). This idea is at the heart of many macroeconomic models, which predict that during normal times, government spending leads to an increase in interest rates, crowding out private investment and lowering future economic output. This occurs as the government spending shock leads to excess demand for resources that may be offset by an increase in interest rates to induce households to reduce consumption and firms to reduce investment, allowing the market to clear the disequilibrium (Devereaux et al., 1996; Murphy & Walsh, 2022). Many empirical works showed that government spending does not have effects, or it has a negative impact on interest rates. Murphy and Walsh (2022) presented a review of applied works that estimated the relationship between government spending and credit market. Among the first, a study by the US Treasury Department (1984), by estimating the impact of a government deficit shock on the real interest rates, finds negative and statistically significant coefficients or positive but insignificant coefficients, according to different specifications of the model. Barro (1987), exploiting the military spending data for the United Kingdom from 1700 to the end of the First World War, finds a positive impact of government spending on real interest rates in the long run only. The evidence of the study by Evans (1987), exploiting a dataset for the United States with monthly data from June 1908 to March 1984, points at a negative effect of current and past government spending on the commercial paper rate, or on Moody's AAA bond rate, or on the *ex post* real commercial paper rate. More recent studies, tackling the issue of endogeneity bias, do not find evidence that identified exogenous shocks to government spending lead to higher interest rates. Edelberg et al. (1999), using a VAR model, find evidence of a negative response (in the short run only) of three different real interest-rates (using 3-month, 1-year, and 2-year Treasury bill yields) to a government spending shock identified through a narrative approach, exploiting the Ramey and Shapiro (1997) episodes. Eichenbaum and Fisher (2005) also use a narrative approach identifying scheme by extending the Ramey–Shapiro episodes with the addition of the 9/11 terror attack. The response of the real rate on Moody's BAA corporate bonds (with an average maturity of 20 years) to an exogenous government spending shock is negative for the first three quarters, whereas the subsequent positive response is statistically insignificant. Mountford and Uhlig (2009) implement a vector autoregression (VAR) model, involving data for the US economy from 1955 to 2000. They combine sign restrictions and zero restrictions to identify the shocks and find that a government spending shock does not lead to an increase in interest rates. Fisher and Peters (2010) propose a new identification strategy, exploiting data on excess stock returns of major US Department of Defense contractors. As a measure for interest rates, they consider the log of the nominal gross three-month treasury bill rate. Consistent with the results discussed above, exogenous government spending is not associated with a substantial change in interest rates. Ramey (2011) implements a narrative approach to identify government spending shocks and uses the three-month Treasury bill rate and the real interest rate on BAA bonds as a measure for interest rates. The results show that the former falls slightly after a positive government spending shock, but the response is not statistically different from zero, while the latter falls significantly for one year and then rises and exceeds 0, before falling again. Corsetti et al. (2012) implement two identification strategies. The first one follows Blanchard and Perotti (2002), whereas the second one exploits fiscal policy changes related to wars and military build-up as in Ramey and Shapiro (1997). The results for the long-term interest rate show that it increases after a government spending shock, but this increase is not statistically significant, and falls afterwards. D'Allesandro et al. (2019) use a structural vector autoregression (SVAR), and their results show that the real interest rate falls after a positive government spending shock. Finally, Murphy and Walsh (2022) show, for the US, a decrease in the Treasury's General Account to a one standard deviation government spending shock identified by an SVAR with real government spending, real tax receipts, and log real GDP. According to Murphy and Walsh (2022) these findings show that US government finances part of its spending using money-like assets, implying an excess supply of loans that leads to a reduction in long-term interest rates.

In Europe, the link between fiscal policy and interest rates has not been extensively studied, unlike in the US (Faini, 2006). Some authors have focused on the impact of fiscal policy on government spreads or sovereign bond interest rates, finding that a positive fiscal shock leads to an increase in real sovereign interest rates (Afonso & Strauch, 2003; Bernoth et al., 2003; Burriel et al., 2009; Codogno et al., 2003). As for Italy, to the best of our knowledge there are no specific studies that aim at estimating the impact of fiscal policy on interest rates and credit markets and the only exception are studies including interest rates in a VAR to estimate the effects of fiscal policy on the real economic activity. For instance, Giordano et al. (2007) estimate the effect of fiscal policy on the Italian economy using the VAR model and find that a positive government spending shock lowers the long-term interest rate on impact and then there is a positive response which is not statistically significant. Therefore, our study aims to fill this gap and provide some specific evidence on the effect of government spending on the credit market in Italy.

3 | EMPIRICAL ANALYSIS

3.1 | Data

We merge data from Bank of Italy's BDS database and the Annual Regional Database of the European Commission's Directorate-General for Regional and Urban Policy (ARDECO). In particular, data for the credit market comes from the Bank of Italy's database, which contains information on the volume of credit that banks grant to different categories of borrowers. We consider loans to firms with less than 20 employees and to producer households, loans to non-financial corporations and producer households and loans to consumer households. To have an overall measure of the credit provided to all the categories of borrowers we sum up the credit volume granted to non-financial corporations and producer households with that granted to consumer households. This dataset consists of a panel of 106 Italian (NUTS 3) provinces at a quarterly frequency from 2011 to 2018.[2] In order to share a common frequency dataset with the ARDECO database (available at yearly frequency), we sum up the volume of loans that firms and households receive during four quarters in each year to obtain annual observations. Table 1 and Figure 1 provide descriptive statistics and the boxplot of the distribution of loans across NUTS 1 regions, respectively, that give

[2]We are aware of the short time series dimension of the panel limiting the empirical analysis to the period during and after the sovereign debt crisis. However, in order to investigate the risk premium channel as the main transmission mechanism from government spending to credit (as suggested by Auerbach et al., 2020) we need to retrieve disaggregated data for different categories of borrowers (especially small businesses and those having access to credit from banks headquartered in the Centre-North and in the Mezzogiorno). This category of data, at NUTS 3 level, is available from the Bank of Italy since 2011 only.

TABLE 1 Descriptive statistics on loan volumes by NUTS 1 regions (euro *per capita*).

	Total		NFCs		Small firms		Households	
	mean	st. dev.	mean	st. dev.	mean	st. dev	mean	st. dev
Centre	108.33	5.65	68.79	6.57	11.12	1.12	39.54	1.26
Islands	47.26	1.13	24.05	1.77	7.03	0.50	23.21	1.13
Northeast	120.10	6.45	82.26	7.02	17.56	1.89	37.84	0.77
Northwest	121.72	5.11	80.24	5.92	12.65	1.20	41.48	1.23
South	48.47	1.11	26.35	1.78	6.52	0.52	22.12	1.09

evidence of a divide between the macro-regions belonging to the Centre–North and to the Mezzogiorno in terms of credit allocation, given a higher share of total loans *per capita* to small firms, Non-Financial Corporations (NFC) and household allocated to the former.

Data on government expenditure, real economy and population come from the ARDECO database. As a measure of real economic activity, we consider GDP at constant prices, for which 2015 is the base year. We use population data to calculate the variables in *per capita* terms. In order to obtain NUTS 3 data for government spending we follow the suggestions of Gabriel et al. (2020) and Brueckner et al. (2022) and we use the gross value added (GVA) of the non-market sector as a proxy for the final consumption expenditure of the General Government (since the latter is the main actor in the non-market sector in Europe, especially in Italy). The GVA of the non-market sector encompasses compensation of employees, including social contributions, consumptions of fixed capital, that measures the reduction in the value of fixed assets due to obsolescence, normal wear and tear, and other taxes minus subsidies (these taxes refer to net taxes on production and they do not include consumption nor corporate taxes). ARDECO data involve government consumption in the following sub-sectors: (i) public administration and defence; (ii) education; (iii) human health and social work; (iv) arts, entertainment, and recreation; (v) other service activities; and (vi) activities of households and extra-territorial organizations and bodies. As pointed out by Gabriel et al. (2020) and Brueckner et al. (2022), the GVA of the non-market sector does not include intermediate consumption of the General Government and only the first three sub-sectors (which cover most of the GVA of the non-market sector) are closely linked to the General Government in the national account. Overall, the GVA of the non-market sector accounts for 70% of the General Government consumption expenditure.[3] Table 2 and Figure 2 provide descriptive statistics and the boxplot of *per capita* government spending across NUTS 1 regions, with the largest share allocated to the Centre and the Islands.

To assess the validity of our proxy, we conduct a graphical and quantitative analysis. We compare the final consumption expenditure of the General Government (FCE) from the AMECO database with the GVA of the non-market sector at the NUTS 0 level from the ARDECO database. Figure 3 shows the two series in log form, and Figure 4 shows the two series of the first difference of these two variables.

We can notice that the two series are quite close to each other, especially after 2000. Indeed, we find that for Italy, the share of the GVA of the non-market sector in the final consumption expenditure of the General Government is about 0.7 only in 1995, 1996 and 1997, it is about 0.85 in 1998 and from 1999 to 2018 is above 0.9. This leads to a significant gap in the first difference from 1995 to 2000, and, as we will show below, it reduces the correlation coefficient and the slope coefficient of the regression of the first difference of the FCE on the GVA.

Therefore, we also conduct a quantitative analysis and compute the correlation coefficient and estimate regressions involving these two variables.

[3]To recover the intermediate consumption, one could use data from the PBL EUREGIO database, which contains information on the intermediate consumption of the non-market sector. However, they are only available at the NUTS 2 level and from 2000 to 2010. Moreover, Gabriel et al. (2020) show that the intermediate consumption account for 30% in the non-market sector and 27% in the general government expenditure and this share is stable over time. More specifically, the authors compute an average standard deviation of 0.018 for the time-varying intermediate consumption share.

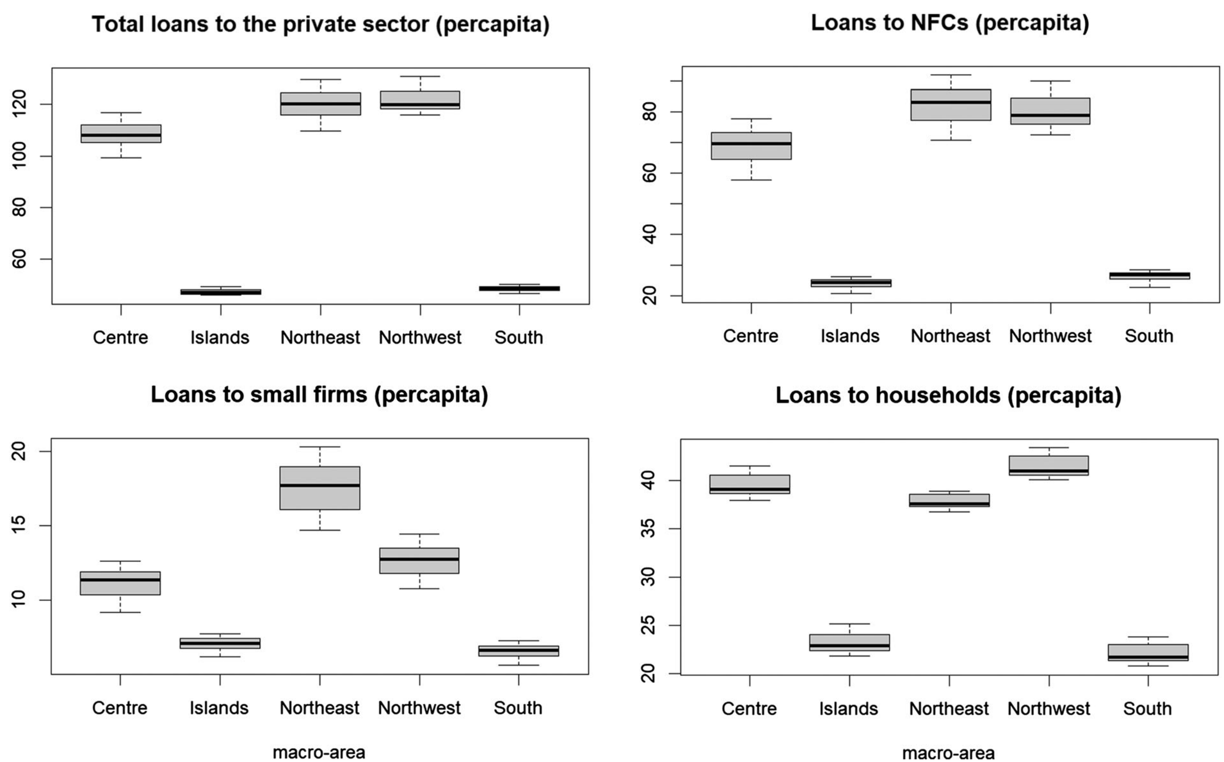

FIGURE 1 Boxplots of loan distribution across NUTS-1 regions (euro *per capita*).

TABLE 2 Descriptive statistics of government spending proxy and GDP by NUTS-1 regions (millions of euro *per capita*, real terms).

	Gov. spend. Proxy		GDP	
	mean	st. dev.	mean	st. dev.
Centre	6306.31	145.37	30335.22	797.62
Islands	5243.80	128.58	18009.71	403.26
Northeast	5246.39	55.68	32804.77	781.80
Northwest	4968.23	68.90	34168.83	865.31
South	4723.96	117.16	18461.45	349.21

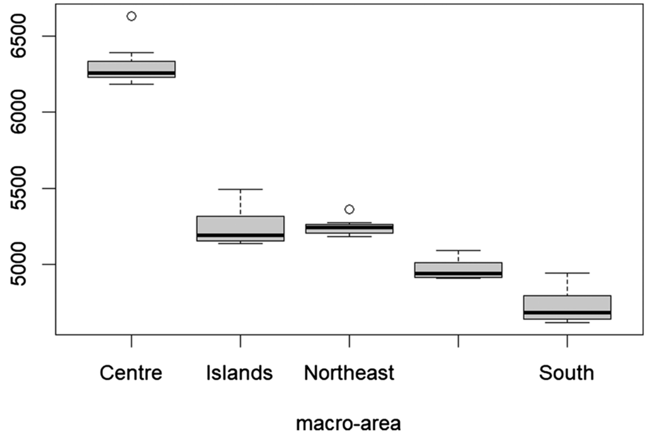

FIGURE 2 Boxplot of the distribution of the government spending proxy in NUTS -1 regions (millions of euro *per capita*, real terms).

Table 3 shows that the correlation coefficients between the final consumption expenditure of the General Government and our proxy, in logarithm, is very close to 1 and it is statistically different from zero. The correlation coefficient between the first differences of these two variables is about 0.77 and statistically different from zero. Table 4 presents the results of the regression of the logarithm of FCE on the logarithm of our proxy and the regression of the first differences of FCE on the first differences of our proxy. The coefficients are close to 1 and 0.8 and statistically different from zero. The standard errors are calculated using the Newey–West estimator of the variance–covariance matrix of the residuals, which allows to control for heteroscedasticity and serial correlation. Thus, the GVA of the non-market sector seems to be a good proxy for government spending at the provincial level. Finally, it is worth noting that this measure does not include investment expenditure, and thus it can be considered a measure of government consumption. Furthermore, it does not include social transfers, and this may help in the identification strategy that we implement (see subsection 3.3).

3.2 | Empirical methodology

To estimate the response of the credit market to a government spending shock we follow the single equation panel regression approach by Auerbach et al. (2020). While the authors focus on the response over a given horizon, we are

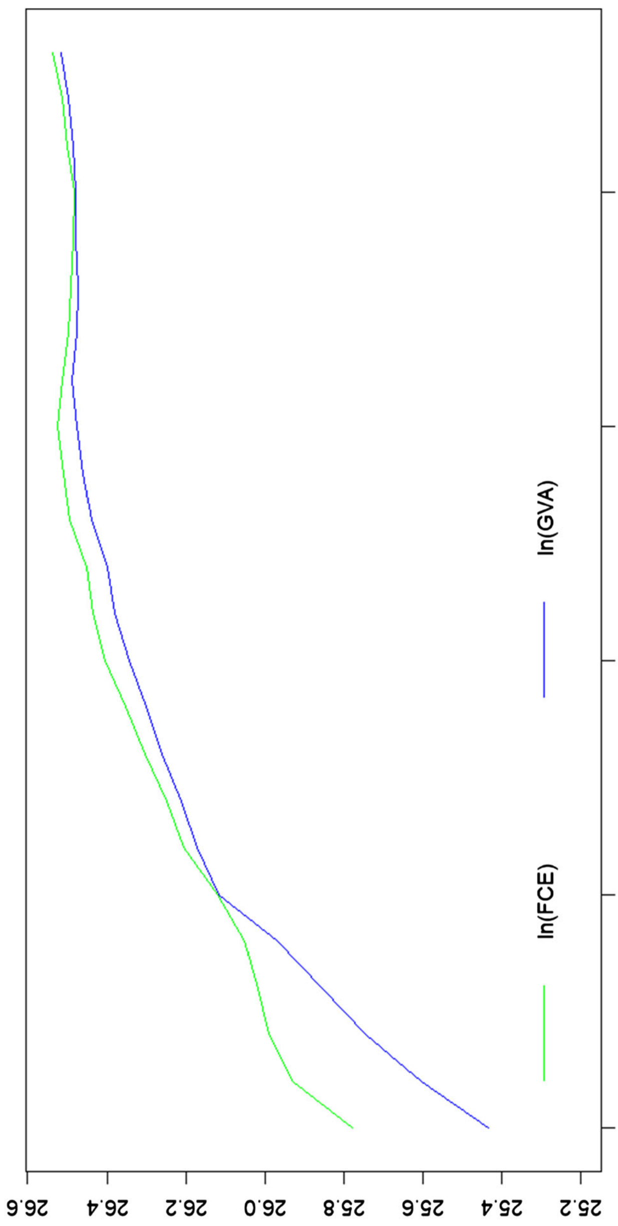

FIGURE 3 Plot of the time series of the final consumption expenditure of the General Government and the GVA of the non-market sector in logarithm.

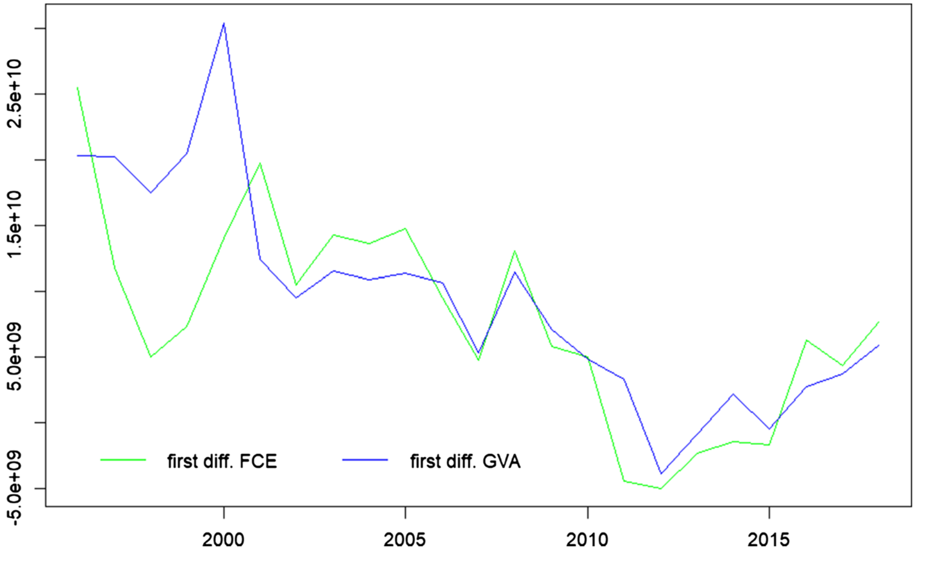

FIGURE 4 Plot of the time series of the first difference of the final consumption expenditure of the General Government and the GVA of the non-market sector.

TABLE 3 Pearson correlation between government spending proxies from AMECO and ARDECO.

Corr (lnFCE, lnGVA)	*Corr (dFCE, dGVA)*
0.9838***	0.7115***

Notes: FCE and GVA (non-market sector) are national government spending proxies from AMECO and ARDECO sources, respectively. The test statistics, based on *z* Fisher Transform, has a *t*-distribution with n-2 d.g.f under null hypothesis of two independent normal distributions.
****p* value < 0.001.

TABLE 4 Regression analysis for Government Spending proxies from AMECO and ARDECO.

	lnFCE
lnGVA	1.0028*** (0.0014)
	dFCE
dGVA	0.7557*** (0.1547)

Newey-West HAC robust standard errors in brackets.
Notes: FCE and GVA (non-market sector) are national government spending proxies from AMECO and ARDECO sources, respectively.
****p*-value < 0.001.

interested in the impulse response profile over several horizons and, for this purpose, we use the Local Projections approach (Jordà, 2005).

Therefore, we estimate the following single equation (for different forecast horizon h):

$$\frac{L_{i,t+h}-L_{i,t+h-1}}{Y_{i,t-1}}=\alpha_{i,h}+\gamma_{t,h}+\beta_h\frac{G_{i,t}-G_{i,t-1}}{Y_{i,t-1}}+\delta_h X_{i,t}+\varepsilon_{i,t+h}, \tag{1}$$

where $h=0,1,2,3$.

The dependent variable is the first difference of the volume of loans $L_{i,t}$ normalized by the lagged value of real GDP $Y_{i,t-1}$. The main explanatory variable is the first difference of real government spending $G_{i,t}$ normalized by the lagged value of real GDP. $X_{i,t}$ is a control variable and, in line with Auerbach et al. (2020), we use one lag of real GDP growth as a measure for the real economic activity, $\alpha_{i,h}$ are provincial fixed effects, and $\gamma_{t,h}$ are time fixed effects (all variables are expressed in *per capita* terms). Since the error term in the local projections follows a moving average process, $MA(h-1)$, we conduct inference robust to heteroscedasticity and autocorrelation using an HAC estimator to calculate the standard errors. In line with Furceri et al. (2021), Gabriel et al. (2020) and Auerbach et al. (2020), we use the Driscoll and Kraay (1998) estimator, which not only controls for heteroscedasticity and serial autocorrelation, but also for cross-sectional dependence across units.

The time span involves the interval 2011–2018 and the inclusion of time fixed effects allows us to control for common shock such as ECB monetary policy interventions (see Gabriel et al., 2020; Nakamura & Steinsson, 2014). The inclusion of provincial fixed effects allows us to control for unobserved heterogeneity across Italian provinces, in order to capture the presence of significant territorial differences in Italy. The impulse response coefficient of interest is β_h and it is estimated by OLS. Since the short data span (due to availability of the loans to small firms dataset) involves only 8 years, we choose to estimate the response up to 3 years and the coefficient β_0 will be interpreted as the impact multiplier, whereas β_h and for $h=1,2,3$ will measure the response of the endogenous variable in $t+h$ to a shock to public spending in t.

3.3 | Identification strategy

Given the small T, large N feature of the panel dataset used we cannot rely on the identification schemes implemented in a SVAR framework. Consequently, to address endogeneity issues, we implement an identification strategy developed by Bartik (1991) that relies on the use of so-called "shift-share instruments" in a panel data regression by interacting a time-invariant variable that varies across cross-sectional units with a time-varying factor which is constant across cross-sectional units. In line with Gabriel et al. (2020) (who focus on European government spending data) we construct an instrumental variable for government spending by, first, constructing the time invariant share:[4]

$$s_i = \frac{\overline{G_i}}{\overline{G_{ITA}}}, \tag{2}$$

that is the ratio of the average government spending in province i over the full sample for which the dataset for public spending is available (1980–2018), to the average national government spending over the same period.

If the ratio is greater than one, then it means that on average, local unit i receives more public sources *per capita* than the national average. The time invariant share measures the exposure of local unit i to common shocks to national public spending. Second, the other interaction term is the time-varying common factor, that is the annual change of national government spending normalized by the lagged value of real GDP:

$$g_t = \frac{G_{ITA,t} - G_{ITA,t-1}}{Y_{ITA,t-1}}. \tag{3}$$

Finally, the instrument is constructed by computing the Kronecker product between the vector containing the shares and the vector containing changes in national government spending, which corresponds to the multiplication of the share of each province by the changes in national expenditure in each year:

[4]See also Nakamura and Steinsson (2014), Auerbach et al. (2020) for the use of Bartik (1991) instruments to identify US government spending shocks.

$$\begin{array}{c}\boldsymbol{Bartik} = \boldsymbol{s} \otimes \boldsymbol{g} \\ Bartik_{i,t} = s_i \times g_t\end{array} . \tag{4}$$

The idea of the identification scheme implemented here is that changes in national government spending should be exogenous to local economic conditions, when especially the level of territorial disaggregation is high enough to believe that no local unit is economically and politically important. This assumption is stronger in the case of disaggregation at the NUTS 3 level since it is hard to say that a specific province can directly influence the decision of the central government. However, the main problem here could be the vector of shares, since they could be related to local economic conditions, namely, local units that are facing a negative phase of the cycle or lower long run growth compared to other local units, receive more government spending relative to the national average and therefore they would have a greater value of s_i. Thus, the source of endogeneity, that is, the violation of the identifying assumption, could come from the vector containing the shares. However, following Gabriel et al. (2020) and also Nakamura and Steinsson (2014), we test for the shares' endogeneity in the following subsection.

3.3.1 | Test for the exogeneity of the instrument

First, we check whether the shares are sensitive to local business cycle, proxied by the standard deviation of each local unit's real GDP growth, and compare it to the shares (see Nakamura & Steinsson, 2014). First, we note that the standard deviation of each local unit's real GDP growth does not change substantially across units and, in particular,

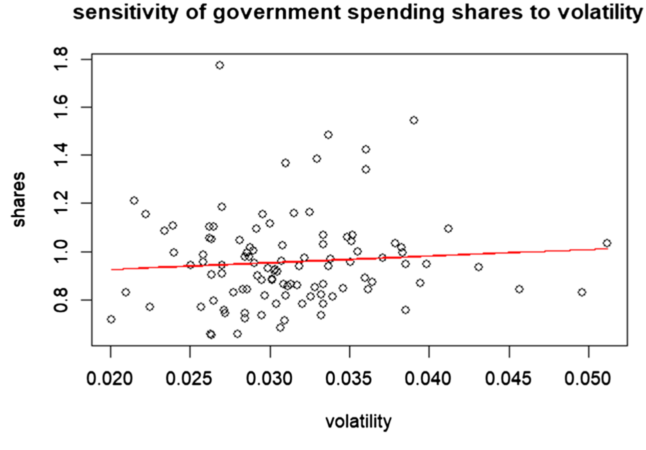

FIGURE 5 Scatterplot of shares and volatility of real GDP growth. *Note*: The red line is the estimated regression line.

TABLE 5 Test for the exogeneity of the instrument a).

	GVA shares	
	Estimate	Std. errors
Intercept	0.8702***	0.1085
GDP volatility	2.7589	3.4161

Notes: The results in the table show test for the sensitivity of shares to GDP volatility (the BreuschPagan test does not reveal the presence of heteroscedasticity; we apply the OLS estimator for std. errors).
***p-value < 0.001.

TABLE 6 Test for the exogeneity of the instrument b).

	Bartik instrument	
	Pooled OLS	Within FE
Intercept	0.003*** (0.0001)	
RSBC	0.003 (0.003)	0.002 (0.002)
time FE	no	yes
Individual FE	no	yes

Standard errors in brackets. Those for FE estimator are Driscoll and Kraay (1998) adjusted standard errors.
Notes: The results in the table show the sensitivity of the Bartik instrument to the measure of the relative stance of the business cycle (RSBC).
***indicates statistical significance at 0.1% level.

it shows similar values for local units with shares above and below the median (see Table A1 in the Appendix). Furthermore, we construct a scatter plot relating the time-invariant shares to the standard deviations of GDP growth for each province and interpolate the scatter by estimating the linear regression line, using the shares as the dependent variable (Figure 5). A positive and statistically significant coefficient would invalidate the identification scheme, since it would suggest that the higher the volatility of the local economy, the higher the share of government spending. The number of provinces is 106, hence we have 106 observations. We can see that there is no obvious relationship between the shares and volatility. The results of the regression analysis are in Table 5. The coefficient associated with GDP volatility is positive but statistically insignificant, thus it appears that the shares calculated on our proxy for government spending are not sensitive to local economic volatility.

Second, following closely Gabriel et al. (2020), we check whether regions that become poorer relative to other regions receive more public spending. If this was true, then the identification hypothesis would be violated. To do this, we construct a measure of the relative stance of the business cycle as the difference between the annual GDP growth rate of each province and the average annual growth rate of all other provinces. We regress the national government spending growth interacted with the shares, that is our instrument, on this indicator of the relative stance of the business cycle. If the coefficients were negative and statistically significant, it would mean that national public spending would increase when local units with a larger share, that is, regions receiving a larger volume of public spending, become poorer than other regions.

Table 6 shows that the coefficient is positive but very close to zero and is statistically insignificant, even controlling for time and unit fixed effects. Furthermore, as pointed out in subsection 3.1, our proxy for government spending does not include social transfers, that are a cyclical component of government spending, and this may also help in our identification assumption. Thus, we conclude that our identification assumption, relying on the exogeneity of the instruments, is valid.[5]

3.3.2 | Tests for the relevance of the instrument

So far, we have discussed only the exogeneity assumption and tried to give some evidence in favour of it. However, another important condition of the instrumental variable that must be satisfied is the relevance, meaning that the

[5]The ARDECO and ISTAT dataset allow to retrieve data for current and capital account government spending only at NUTS 2 level. Another source, only available at NUTS 2 level, is the database Spesa Statale Regionalizzata of the General Accounting Office (Ragioneria Generale dello Stato) at the Italian Ministry of Economy and Finance, which provides spending of the various Italian departments. The focus on NUTS 3 is motivated by the use of a Bartik instrument which does not suffer from violation of exogeneity assumption, as in the case of NUTS 2 data (results available upon request).

TABLE 7 First stage regression and test for the relevance of the instrument.

	Government consumption (GVA)	
	Estimate	Std. errors
Bartik	0.7267***	0.0846
F-statistic	39.9898	
*** p-value < 0.001		

Estimator	F-statistic
White (1980)	27.0546
White (1984)	79.1388
Arellano (1987)	15.6025
Driscoll and Kraay (1998)	13.6331
Kleibergen-Paap	26.06

instrument must be highly correlated with the endogenous variable. More specifically, the relevance condition can be written into the formula as follows:

$$E\left[\left(\frac{G_{i,t}-G_{i,t-1}}{Y_{i,t-1}}\right)\times Bartik_{i,t}|\alpha_i,\gamma_t,X_{i,t}\right]\neq 0. \tag{5}$$

Here we perform several tests to provide evidence of the relevance of our instrumental variables.

First, we present the results of the first stage regressions. The Bartik instrument is a special type of instrumental variable that tries to isolate an exogenous part of the endogenous variable by decomposing the latter into different dimensions and trying to exploit the exogenous components. Thus, by construction we expect the relevance condition to be satisfied for this type of instrument. We also expect the coefficient of the first stage to be positive, but it should not equal one (Breuer, 2022). To test for the relevance of our instrument, we can run the first-stage regression and test the significance of the first-stage coefficient using the *F*-test. We compute the *F*-statistic using different estimators of the variance–covariance matrix. Thus, in addition to the usual *F*-test, we calculate it using: (i) White's (1980) correction for overall heteroscedasticity but without serial correlation; (ii) White's (1984) correction but assuming constant variance within groups; (iii) Arellano's (1987) estimator to control for both heteroscedasticity and serial correlation; and (iv) Driscoll and Kraay's (1998) estimator to control for heteroscedasticity, serial correlation, and correlation across units. We compare the *F*-statistics with the rule of thumb of Staiger and Stock (1997), which suggests rejecting the hypothesis that the instrument is weak if the *F*-statistic is greater than 10. In addition, following Brueckner et al. (2022) and Furceri et al. (2021), we compute the Kleibergen-Paap rk Wald *F*-statistic to conduct further tests on our instrument. Andrews et al. (2019) show that the Kleibergen-Paap rk Wald *F*-statistic is equivalent to a non-homoscedasticity-robust *F*-statistic to test the relevance of the first-stage coefficient, in the case of one endogenous regressor and one instrument and must be compared to the critical values of Stock and Yogo (2005). If the Kleibergen-Paap statistic is greater than these critical values, then we can reject the hypothesis that our instrument is weak. Table 7 shows the results of the first stage regression and the *F*-tests. The results show that the first-stage coefficient associated with the Bartik instrument is positive as expected and statistically significant. The *F*-statistics are all well above the threshold value of 10 and the Kleibergen-Paap statistic is greater than the critical values of Stock and Yogo (2005).[6] Therefore, our instrument is relevant.

[6]The Stock and Yogo (2005) critical values are as follows: (i) for 10% maximal IV size the critical value is 16.38; (ii) for 15% maximal IV size is 8.96; (iii) for 20% maximal IV size is 6.66; and (iv) for 25% maximal IV size is 5.53. These are the critical values of the size method, in which a researcher control for the size of the Wald test of the null hypothesis that the coefficient is equal to zero. This method suggests rejecting the hypothesis that the instrument is weak if the *F*-statistic of the first stage is greater than these critical values.

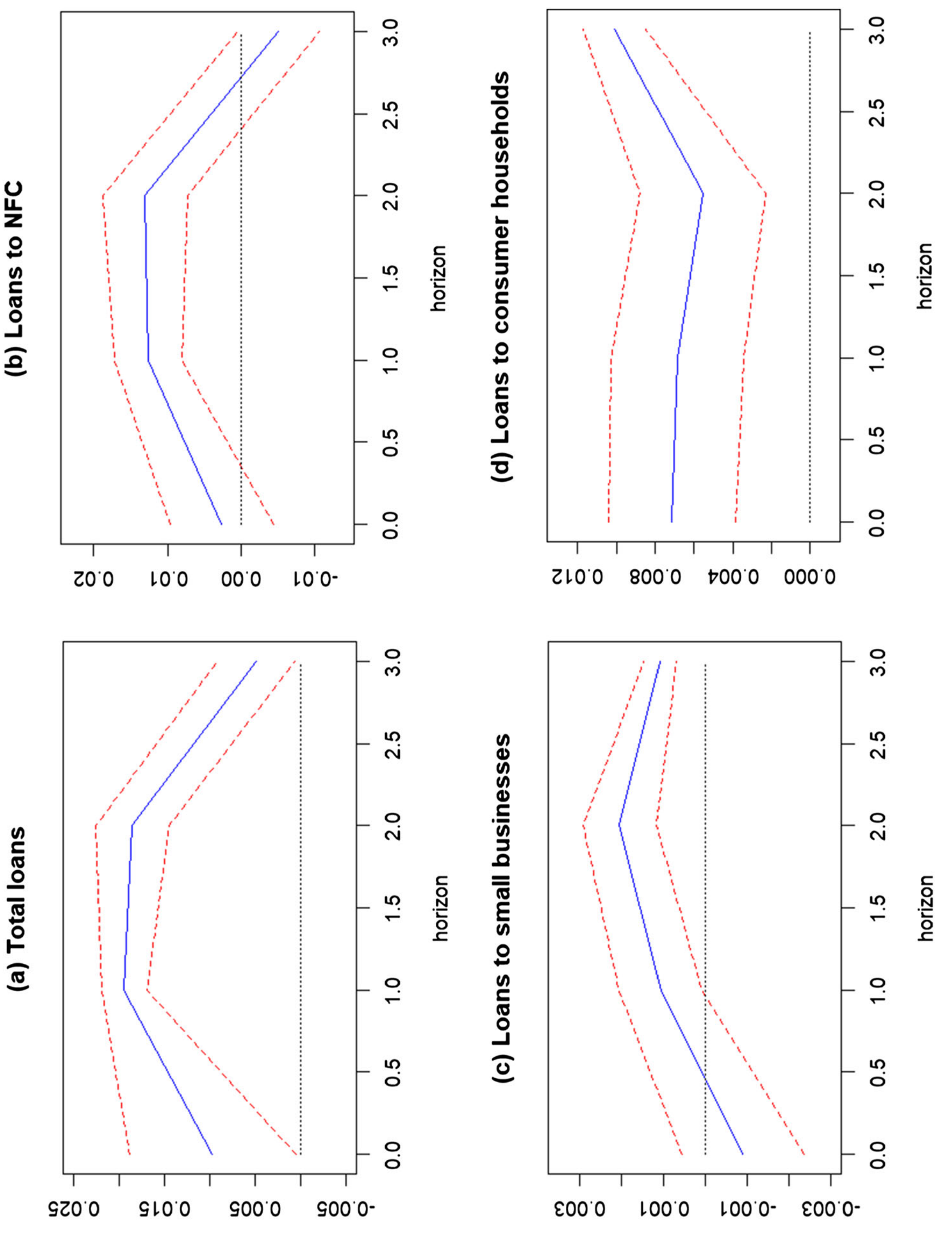

FIGURE 6 Impulse response functions.

3.4 | Local projection IV and size bias

In the first stage of the IV local projection approach, we estimate (using the whole sample data for the GVA of the non-market sector observed for 106 Italian (NUTS 3) provinces, over the time span running from 1980 to 2018) the following panel regression model:

$$\frac{G_{i,t}-G_{i,t-1}}{Y_{i,t-1}}=\alpha_i+\gamma_t+\beta Bartik_{i,t}+\delta X_{i,t}+\varepsilon_{i,t}, \tag{6}$$

where the dependent variable is the province specific annual change of government consumption normalized by the lagged value of the real GDP and the explanatory variable is $Bartik_{i,t}$ that is the instrument described above, by computing the interaction between the time invariant share and the time varying common factor proxied by annual change in national government consumption normalized by the lagged value of the real GDP. Moreover, we control for a lag of the real GDP growth and for time and provincial fixed effects.

In the second stage of the Local Projection IV, we collect the fitted values from the first stage regression, that is $G_{i,t}-\widehat{G_{i,t-1}}/Y_{i,t-1}$, only for the sub-sample 2011–2018 to match the availability of credit market data, and we estimate local projections regression equation:

$$\frac{L_{i,t+h}-L_{i,t+h-1}}{Y_{i,t-1}}=\alpha_{i,h}+\gamma_{t,h}+\beta_h\frac{G_{i,t}-\widehat{G_{i,t-1}}}{Y_{i,t-1}}+\delta_h X_{i,t}+\varepsilon_{i,t+h,} \tag{7}$$

where $h=0,1,2,3$.

Figure 6 shows the Impulse Response Functions estimated by Equation (7) (Table A2 in the Appendix shows the results). It is quite clear that a government consumption shock stimulates the credit market.

The largest (and statistically significant) reaction of a change in total loans (relative to GDP) to a 1% increase in government consumption (relative to GDP) is nearly equal to 0.02% and it is observed one and two years after the shock. The NFC sector is the one contributing the most of the (mild) growth in the volume of loans over the same forecast horizon and it is nearly equal to 0.013%. The impact on the household and on the small business sector is half and a quarter of the one experiencing the NFC sector.

Overall, the positive impact of unanticipated government spending, as "new production" (in line with Auerbach et al., 2020) on credit can be ascribed to the improved bank risk profile assessment of different categories of borrowers in the private sector. These findings are confirmed studying the impact of the unanticipated government spending on a proxy of the risk premium which is the non-performing loan ratio for the whole private sector and for the non-financial corporations and household categories.[7]

In summary, the empirical evidence supports the financial accelerator and the banks liquidity preference channels (highlighted in the Introduction) through which local government spending impact on the volume of loans.[8]

Our findings show that, since that an increase in public spending stimulates credit growth for both larger and, to a smaller extent, for smaller firms, fiscal policy does not benefit the latter category any more than it does for larger firms. As also mentioned in Section 1, this issue is relevant since, as argued by Dow (1996), small businesses face more difficulties to obtain bank loans due to information asymmetries. Our results show that, in absolute terms, an

[7]Specifically, Table 8 shows the results of the estimation of a panel regression similar to Equation (7), having the first difference of the logit transform of the non-performing loan ratio, as the dependent variable. A 1% increase in government spending relative to GDP lowers, on impact, the odd ratio of the probability of default to the probability of solvency of the private sector, of the total of NFCs (including producer households) and of the households, by 0.74%, 0.57%, and 0.12% respectively.

[8]The analysis of the impact of public spending on credit volumes through the channel of bank funding providers liquidity preference (hence related to their cost of funding) would require an investigation of the portfolio rebalancing of depositors and equity investors. We leave this for future research.

TABLE 8 Results for non-performing loan rates.

	Non-performing loan rate (logit transformation)		
	NFCs and producer households	All borrowers excluding financial and monetary inst.	Consumer households, nonprofit organizations and residual values
shock	−0.57* (0.25)	−0.74** (0.25)	−0.12 (0.20)

Notes: Driscoll and Kraay (1998) robust standard errors in brackets.
**, * indicate statistical significance at 1% and 5% levels respectively.

increase in government spending stimulates lending to small firms, although this effect tends to be smaller than the one associated with firms relatively larger.

This result is relevant for the Italian context, given the dominant role of small companies in the Italian economy, and points towards a greater focus on the issue of access to credit for smaller businesses.

3.5 | Local projection IV and home bias

3.5.1 | Home bias

While the empirical evidence of Gabriel et al. (2020) shows lower fiscal multipliers for peripheral European countries than for central European countries, our study focuses on the geographic divide characterizing the credit market in Italy. This is motivated by the bank consolidation process through M&A occurring in the last three decades in Italy. As pointed out by Papi et al. (2015), the banking consolidation process has given banks in the North a central position in the national credit market, relegating banks in the South to a small local market, increasing the so-called "functional distance", that is, the geographical and economic distance between the banks' headquarters, namely, the offices where credit decisions are made, and the bank branches, which are those closest to local communities. (Alessandrini et al., 2009). In turn, increasing functional distance may influence the probability of local borrowers being credit rationed when they are in an area predominantly populated by banks whose headquarters are far away. This is what the literature on "home bias" points out. The greater the functional distance, the more difficult it is to assess and collect "soft and social-embedded information", that is, information that cannot be retrieved only by analysing the balance sheets and financial health indicators of borrowers (the so-called "hard information"). Large banks rely more on hard information, while small local banks establish a closer relationship with small local businesses. Moreover, functional distance increases with bank size. This suggests that larger banks are less inclined to collect soft information. Thus, the bank-firm relationship is stronger when it involves small local banks and small local firms than when it involves large distant banks and small local firms (Berger et al., 2005). Presbitero et al. (2014), exploiting NUTS 3 level data for Italy, find evidence of the presence of a home bias in Italy, and the penetration of distant banks into local credit markets exacerbates the credit crunch in the post-Lehman period.[9]

Therefore, as explained in the Introduction, our third contribution to the literature focusing on the impact of government spending on credit growth, aims at evaluating whether the aforementioned home bias is reduced via fiscal policy shock in a specific macro-region. As argued by Auerbach et al. (2020), public spending could increase credit supply by lowering the risk assessment of local borrowers by lenders, because it stimulates the economy and leads to a reduction in the probability of a local recession, thereby lowering risk premia. Thus, an expansion of government spending may improve the balance sheet of borrowers, thereby influencing the "hard information" that banks consider when making lending decisions. Therefore, it may be interesting to test how a public spending shock in Italy affects the credit market, by separating the credit provided by banks based in the central and northern regions from that provided by banks based in the southern regions.

[9]A recent study from the Bank of Italy (2021), identifying a credit supply diffusion index through the dataset from the Bank Lending Survey, confirms a loan supply contraction (both for firms and households) in the Southern regions relative to the Centre-North over 2011–2013 (post Lehman period).

3.5.2 | Empirical analysis of the geographical location of the borrowers

According to the classification of the Bank of Italy's BDS Database, we consider the following territorial aggregation: (i) the northern regions are Piedmont, Valle d'Aosta, Liguria, Lombardy, Trentino-Alto Adige, Veneto, Friuli Venezia-Giulia, Emilia-Romagna; (ii) regions in the center are Tuscany, Umbria, Marche, Lazio; and (iii) the southern regions are Abruzzo, Molise, Campania, Apulia, Basilicata, Calabria, Sicily, Sardinia.

We divide our sample into two, one for the North–Central and one for the Southern regions and estimate the local projections (7) to obtain the impulse response functions for each area. The IRFs are represented in Figure 7, whereas the results are in Table A3 in the Appendix.

Figure 7(a) shows that the positive effect of a shock to the growth of government consumption relative to GDP on the total loans growth relative to GDP is associated only with the Centre–North, producing (by taking the sum of the statistically significant impulse response coefficients across the different forecast horizon) an overall cumulative impact of 0.08% after three years. On the contrary, the response of the total loan growth (relative to GDP) to a 1% increase in the government consumption (relative to GDP) is negative for the Southern regions.

These findings are consistent across the different sectors. More specifically, the largest positive contribution in the Centre–North can be attributed to the NFC sector (showing an overall cumulative impact equal to 0.04% by taking the sum of the impulse response coefficients for 1, 2 and 3 years ahead), while the household and small business sector exhibit a milder positive impact, given the associated cumulative response equal to 0.0392 and 0.0052%, respectively. The largest negative response of total loan growth in the Southern regions is recorded for the NFC and the household sector for which we observe a fall by nearly 0.08%, whereas the impact on small businesses is negligible (see Figure 7, panel (b), (c) and (d)).

3.5.3 | Empirical analysis of the geographical location of the banks

In the Bank of Italy's BDS database, data on loans classified by bank location are available. In particular, it is possible to distinguish loans granted by banks based in central and northern Italy and by Cassa Depositi e Prestiti (CDP) from those granted by banks based in southern Italy. Thus, we estimate local projections (7) for four sub-samples: loans granted by Central–Northern banks and CDP to all borrowers, that is households, firms, and local public administrations either in Central–Northern provinces or in Southern provinces; loans granted by Southern banks to all borrowers, that is households, firms and local public administrations either in Central–Northern province or in Southern provinces. Figure 8 shows the Impulse Response Functions (Table A4 in the Appendix reports the results).

First, while the empirical findings in Figure 7 (see blue lines) show an increase in credit to private sector local borrowers in the Central–Northern provinces due to a public spending shock, the rightmost top panel in Figure 8 shows a negligible impact of this shock on the credit provided by banks headquartered in the Centre–North to the whole set of local borrowers (including public administration) in the Centre–North. These empirical findings would suggest a curb in credit to local public administration in the Centre–North offsetting the increase in the loans granted to the private sector as a response to a government spending shock.[10]

Moreover, the fall in the growth (relative to GDP) of loans to Southern borrowers in response to a public spending shock in the South observed in Figure 8 can be ascribed to a contraction in the credit supply from banks headquartered in the Centre–North which more than offset the credit supply expansion from banks headquartered in the South. In particular, the cumulative response of the growth in the loans from banks headquartered in the Centre North to local borrowers in the South to 1% increase in Southern government consumption relative to GDP is equal to −0.51%. The empirical analysis shows that the only improvement in the risk profile of borrowers in the

[10]Table TDB from the Bank of Italy shows that, among different economic sectors, the largest drop (occurring over 2012–2018) in the annualized rate of change in loans across different macro-regions is the one for government sector especially in the Northern and Southern geographical areas (with values averaging about −5% and −4% respectively).

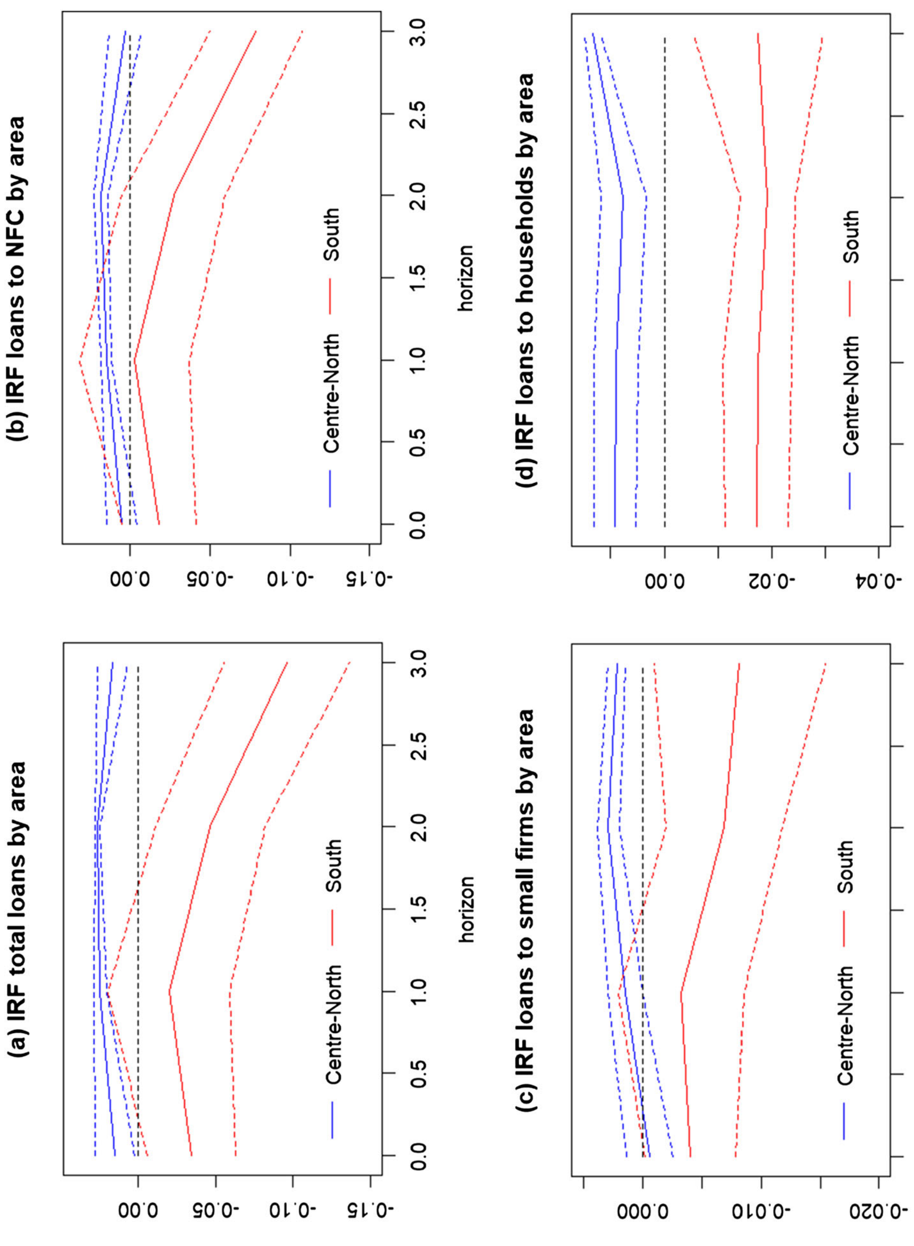

FIGURE 7 Impulse response functions by macro-geographical area.

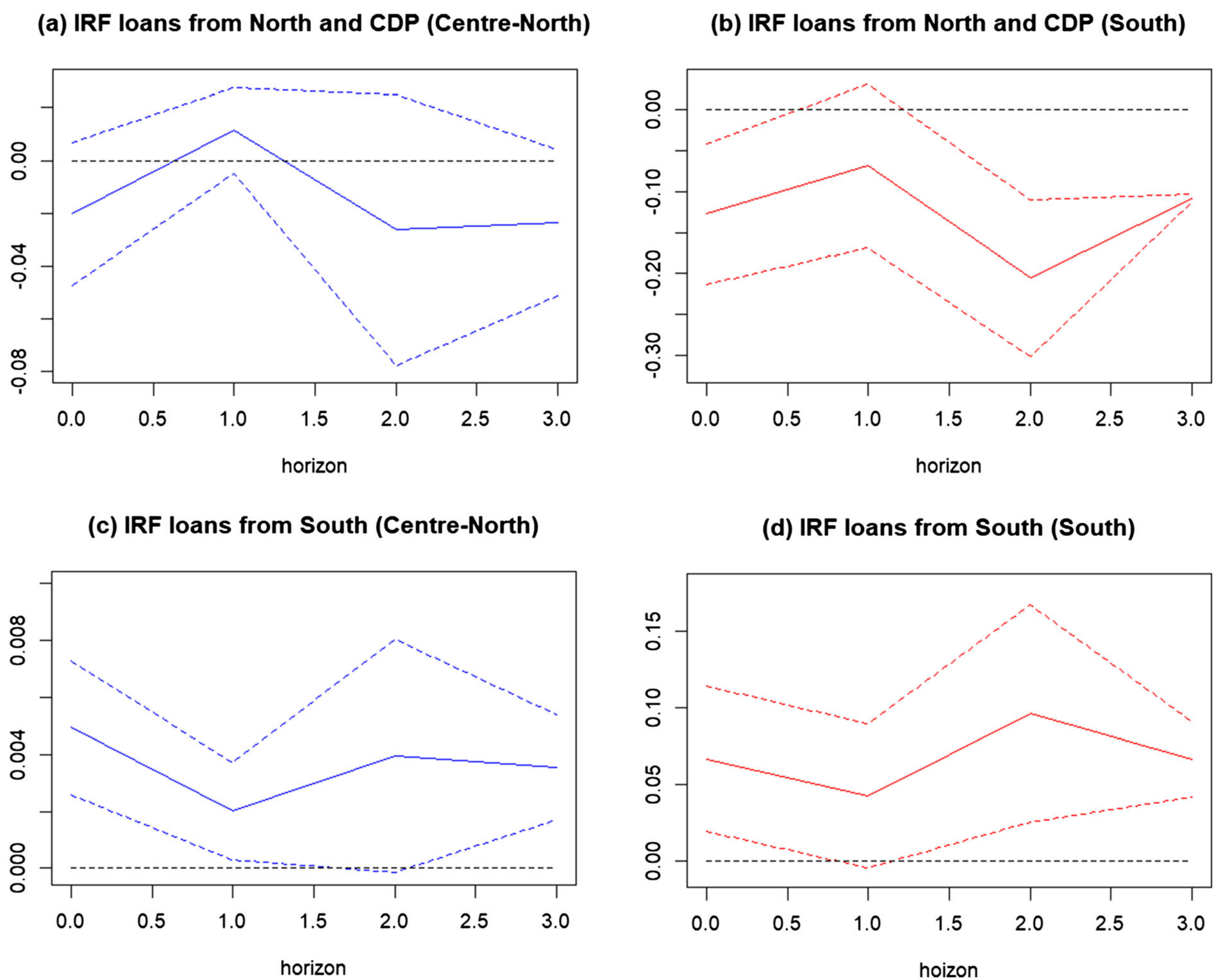

FIGURE 8 Impulse response functions by geographical macro-area and bank headquarters.

Mezzogiorno, assessed by banks with headquarters in the Mezzogiorno, can be associated with an increase in public spending in the same macro area. In particular, the cumulative response of the growth in the loans from banks headquartered in the South to local borrowers in the South to 1% increase in Southern government consumption relative to GDP is equal to 0.27%.[11]

Consequently, the process of bank polarization in the North does not benefit local borrowers in the South, and even an injection of public sources into the local economy does not alleviate this problem. Rather, the credit granted by banks in the Centre–North declines after an increase in local government spending in the South, and this may be related to the issue regarding the economic and institutional divide between the two areas.

4 | ROBUSTNESS ANALYSIS

A robustness check has been carried by normalizing both the change in credit and the one in government spending by total GVA at constant prices instead of real GDP. Thus, we substitute $Y_{i,t}$ with $GVA_{i,t}$ in Equation (7) and estimate the models in section 3.4 and 3.5. The results, shown in Tables A5–A8 in the Appendix, are qualitatively similar, confirming the previous empirical findings.

5 | CONCLUSIONS

In this study we assess the effects of government spending on credit growth employing the local projection approach developed by Jordà (2005). We focus on the Italian economy and exploit a panel dataset of 106 Italian provinces over the period 2011–2018. The identification of the public spending shock is achieved by constructing a Bartik (1991), or "shift-share", instrument.

The empirical evidence shows a mild positive effect of a one percentage point increase in public spending relative to GDP (or GVA) on the growth of loan volume relative to GDP (or GVA). The positive effect involves different categories of the private sector: non-financial corporations, small businesses, and households. This is motivated by first, the risk premium channel transmission mechanism from government spending to credit. As pointed out by Auerbach et al. (2020), unanticipated government spending is interpreted as "new production" having a direct impact on the bank risk profile assessment of different categories of borrowers in the private sector. Moreover, we assess the impact of the identified unanticipated government spending on a proxy of risk premium, which is the non-performing loan ratio for the whole private sector and for the non-financial corporations and household categories. Second, other motivations come from the liquidity preference of banks (see Dow, 1996, and also Palley, 2002, 2017) associated to minimum capital requirements constraining bank lending. Government spending can in this way help reduce this phenomenon, thus limiting the risk of credit cuts or stimulating the issuance of new loans. These results have relevant policy implications, as they provide evidence that in a liquidity trap (which characterizes the sample period we consider), government spending policy can stimulate the credit market together with monetary policy.

However, we observe that government spending does not help to ameliorate the "size bias", that is, the financial constraints of small firms, since they benefit less than the overall category of non-financial firms from increased government spending. This also has relevant implications in the Italian context, as the presence of small or medium-sized enterprises dominates the Italian economy. Moreover, the empirical analysis shows that the only improvement in the risk profile of borrowers in the Mezzogiorno assessed by banks with headquarters in the Mezzogiorno can be associated with an increase in public spending in the same macro-area. These empirical findings show that government spending does not help to ameliorate the home bias in credit related to the process of bank consolidation in Italy.

[11]The evidence of a stronger response of local credit to local public spending can be ascribed to, first, a limited role of government spending in reducing information asymmetries arising in terms of functional distance, and to the credit portfolio rebalancing channel induced by bank liquidity preference (see Dow, 1996; Palley, 2002, 2017).

Together, these results imply that government consumption, although it might be useful on aggregate to revitalize the credit market, is not the only policy tool to dampen credit market territorial differences in Italy. One possibility (which is scope for future research) would be to explore the impact of government investment and/or targeted financial instruments, provided through EU funding, to boost credit (without any crowding out of private credit).

We are aware that the use of data at NUTS 3 level constraints the empirical analysis to the study of the impact of only one category of government spending, namely public consumption. Moreover, the use of NUTS 2 data would enrich the analysis, allowing the distinction between different categories of government spending. However, this extension (which is scope for further research) would require the use of instruments different from Bartik which suffer from violation of the exogeneity assumption once we move to a level of aggregation higher than NUTS 3.

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REFERENCES

Afonso, A., & Strauch, R. (2003). Fiscal policy events and interest rate swap spreads: Some evidence from the EU. *Journal of International Financial Markets, Institutions & Money*, *17*(3), 261–276.

Alessandrini, P., Presbitero, A. F., & Zazzaro, A. (2009). Banks, distances, and firms' financing constraints. *Review of Finance*, *13*(2), 261–307. https://doi.org/10.1093/rof/rfn010

Andrews, I., Stock, J., & Sun, L. (2019). Weak instruments in instrumental variables regression: Theory and practice. *Annual Review of Economics*, *11*(1), 727–753. https://doi.org/10.1146/annurev-economics-080218-025643

Arellano, M. (1987). Computing robust standard errors for within-groups estimators. *Oxford Bulletin of Economics and Statistics*, *49*(4), 431–434.

Auerbach, A. J., Gorodnichenko, Y., & Murphy, D. (2020). Effects of fiscal policy on credit markets. NBER Working Paper 26655.

Bank of Italy. (2021). La domanda e l'offerta di credito a livello territoriale. Economie Regionali n. 43.

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

Bartik, T. J. (1991). *Who benefits from state and local economic Ddevelopment policies?* W.E. Upjohn Institute for Employment Research. https://doi.org/10.17848/9780585223940

Berger, A. N., Miller, N. H., Petersen, M. A., Rajan, R. G., & Stein, J. C. (2005). Does function follow organizational form? Evidence from the lending practices of large and small banks. *Journal of Financial Economics*, *76*(2), 237–269. https://doi.org/10.1016/j.jfineco.2004.06.003

Bernanke, B. S., Gertler, M., & Gilchrist, S. (1999). The financial accelerator in a quantitative business dycle framework. In J. Taylor & M. Woodford (Eds.). *Handbook of macroeconomics* (Vol. 1, part C, pp. 1341–1393). Elsevier.

Bernoth, K., Schuknecht, L., & von Hagen, J. (2003). The determinants of the yield differential in the EU bond market. Working paper Retrieved from http://www.soc.uoc.gr/calendar/ss6/students_papers/Bond-Project1.pdf

Blanchard, O. J., & Perotti, R. (2002). An empirical characterization of the dynamic effects of changes in government spending and taxes on output. *Quarterly Journal of Economics*, *117*(4), 1329–1368. https://doi.org/10.1162/003355302320935043

Breuer, M. (2022). Bartik instruments: An applied introduction. *Journal of Financial Reporting*, *7*(1), 49–67. https://doi.org/10.2139/ssrn.3786229

Brueckner, M., Pappa, E., & Valentinyi, A. (2022). Geographic cross-sectional fiscal spending multipliers and the role of local autonomy: Evidence from European regions. *Journal of Money, Credit and Banking* (forthcoming). https://doi.org/10.1111/jmcb.12974

Burriel, L. P., Castro Fernández, F. D., Garrote Sánchez, D., Gordo Mora, E., & Pérez García, J. J. (2009). Fiscal policy shocks in the euro area and the US: An empirical assessment. Documentos de trabajo/Banco de España, 0930.

Chick, V., & Dow, S. (1988). A post-Keynesian perspective on the relation between banking and regional development. In *Thames Papers in Political Economy* (pp. 1–22). Spring.

Christiano, L., Eichenbaum, M., & Rebelo, S. (2011). When is the government spending multiplier large? *Journal of Political Economy*, *119*(1), 78–121. https://doi.org/10.1086/659312

Codogno, L., Favero, C., & Missale, A. (2003). Yield spreads on EMU government bonds. *Economic Policy*, *18*(37), 503–532. https://doi.org/10.1111/1468-0327.00114_1

Corsetti, G., Meier, A., & Müller, G. J. (2012). Fiscal stimulus with spending reversals. *Review of Economics and Statistics*, *94*(4), 878–895.

D'Allesandro, A., Fella, G., & Melosi, L. (2019). Fiscal stimulus with learning-by-doing. *International Economic Review*, *60*(3), 1–20.

Devereaux, M. B., Head, A. C., & Lapham, B. J. (1996). Monopolistic competition, increasing returns, and the effects of government spending. *Journal of Money, Credit, and Banking*, *28*(2), 233–254.

Dow, S. C. (1996). Horizontalism: A critique. *Cambridge Journal of Economics*, *20*(4), 497–508. https://doi.org/10.1093/oxfordjournals.cje.a013629

Driscoll, J. C., & Kraay, A. C. (1998). Consistent covariance matrix estimation with spatially dependent panel data. *The Review of Economics and Statistics*, *80*(4), 549–560. https://doi.org/10.1162/003465398557825

Edelberg, W., Eichenbaum, M., & Fisher, J. D. M. (1999). Understanding the effects of a shock to government purchases. *Review of Economic Dynamics*, *2*, 166–206.

Eichenbaum, M., & Fisher, J. D. M. (2005). Fiscal policy in the aftermath of 9/11. *Journal of Money, Credit, and Banking*, *37*(1), 1–22.

Evans, P. (1987). Interest rates and expected future budget deficits in the United States. *Journal of Political Economy*, *95*(1), 34–58.

Faini, R. (2006). Fiscal policy and interest rates in Europe. *Economic Policy*, *21*(47), 444–489. https://doi.org/10.1111/j.1468-0327.2006.00163.x

Fisher, J. D. M., & Peters, R. (2010). Using stock returns to identify government spending shocks. *The Economic Journal*, *120*(544), 414–436. https://doi.org/10.1111/j.1468-0297.2010.02355.x

Furceri, D., Ganslmeier, M., & Ostry, J. (2021). *Are climate change policies politically costly? IMF Working Paper*, WP/21/156,. International Monetary Fund.

Gabriel, R. D., Klein, M., & Pessoa, A.S. (2020). The effects of government spending in the Eurozone. *Sveriges Riksbank Working Paper Series*, No. 400, December 2020.

Giordano, R., Momigliano, S., Neri, S., & Perotti, R. (2007). The effects of fiscal policy in Italy: Evidence from a VAR model. *European Journal of Political Economy*, *23*(3), 707–733. https://doi.org/10.1016/j.ejpoleco.2006.10.005

Jordà, Ò. (2005). Estimation and inference of impulse responses by local projections. *American Economic Review*, *95*(1), 161–182.

Mountford, A., & Uhlig, H. (2009). What are the effects of fiscal policy shocks? *Journal of Applied Econometrics*, *24*, 960–992.

Murphy, D., & Walsh, K. J. (2022). Government spending and interest rates. *Journal of International Money and Finance*, *123*, 102598. https://doi.org/10.1016/j.jimonfin.2022.102598

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

Palley, T. I. (2002). Endogenous money: What it is and why it matters. *Metroeconomica*, *53*(2), 152–180. https://doi.org/10.1111/1467-999X.00138

Palley, T. I. (2017). The theory of endogenous money and the LM schedule: prelude to a reconstruction of ISLM. *Brazilian Journal of Political Economy*, *37*, 3–22. https://doi.org/10.1590/0101-31572016v37n01a01

Papi, L., Sarno, E., & Zazzaro A. (2015). The geographical network of bank organizations: Issues and evidence for Italy. Working Paper No. 105. Mo.Fi.R.

Presbitero, A. F., Udell, G. F., & Zazzaro, A. (2014). The home bias and the credit crunch: A regional perspective. *Journal of Money, Credit and Banking*, *46*(s1), 53–85. https://doi.org/10.1111/jmcb.12078

Ramey, V. A. (2011). Identifying government spending shocks: It's all in the timing. *Quarterly Journal of Economics*, *126*(1), 1–50.

Ramey, V. A., & Shapiro, M. D. (1997). Costly capital reallocation and the effects of government spending. Carnegie-Rochester Conference Series on Public Policy 48 (1998) 145-194, North-Holland.

Staiger, D., & Stock, J. H. (1997). Instrumental variables regression with weak instruments. *Econometrica*, *65*(3), 557–586. https://doi.org/10.2307/2171753

Stock, J. H., & Yogo, M. (2005). In D. W. K. Andrews & J. H. Stock (Eds.). *Identification and Inference in Econometric Models: Essays in Honor of Thomas J. Rothenberg*. Cambridge University Press.

US Treasury. (1984). The effect of deficits on prices of financial assets: Theory and evidence. US Government Printing Office.

White, H. (1980). A heteroscedasticity-consistent covariance matrix estimator and a direct test for heteroskedasticity. *Econometrica*, *48*(4), 817–838. https://doi.org/10.2307/1912934
White, H. (1984). *Asymptotic theory for econometricians*. Academic press.

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APPENDIX

TABLE A1 Comparison of government spending shares and GDP volatility.

Province	Shares	GDP volatility	Province	Shares	GDP volatility
Vicenza	0.652	0.026	Ferrara	0.940	0.032
Bergamo	0.654	0.026	Venezia	0.941	0.027
Lecco	0.655	0.028	Napoli	0.944	0.025
Treviso	0.683	0.031	Rimini	0.944	0.040
Varese	0.712	0.031	Taranto	0.948	0.039
Lucca	0.717	0.020	Frosinone	0.950	0.029
Pistoia	0.720	0.028	Avellino	0.956	0.035
Monza e della Brianza	0.732	0.029	Sondrio	0.956	0.026
Biella	0.733	0.033	Sassari	0.959	0.031
Mantova	0.742	0.027	Reggio di Calabria	0.969	0.034
Brescia	0.744	0.028	Bari	0.970	0.032
Lodi	0.756	0.039	Chieti	0.973	0.037
Fermo	0.757	0.027	Latina	0.975	0.029
Prato	0.768	0.026	Catania	0.976	0.028
Massa-Carrara	0.769	0.022	Salerno	0.984	0.026
Reggio nell'Emilia	0.781	0.030	Padova	0.994	0.029
Asti	0.782	0.033	Parma	0.995	0.024
Como	0.784	0.032	Potenza	0.996	0.038
Arezzo	0.795	0.026	Isernia	0.999	0.035
Lecce	0.813	0.034	Nuoro	1.003	0.029
Barletta-Andria-Trani	0.814	0.033	Pordenone	1.017	0.038
Rovigo	0.816	0.030	Grosseto	1.017	0.029
Caserta	0.818	0.031	Terni	1.025	0.031
Foggia	0.822	0.033	Messina	1.029	0.033
Imperia	0.830	0.050	Siracusa	1.031	0.051
Agrigento	0.831	0.028	Vibo Valentia	1.033	0.038

(Continues)

TABLE A1 (Continued)

Province	Shares	GDP volatility	Province	Shares	GDP volatility
Verona	0.831	0.021	Enna	1.044	0.035
Belluno	0.841	0.036	Viterbo	1.048	0.028
Modena	0.841	0.028	Perugia	1.050	0.026
Macerata	0.842	0.029	Palermo	1.056	0.026
Crotone	0.843	0.046	Gorizia	1.060	0.035
Pavia	0.848	0.035	Ragusa	1.066	0.035
Brindisi	0.849	0.033	Rieti	1.069	0.033
Pesaro e Urbino	0.856	0.031	Livorno	1.084	0.023
Alessandria	0.861	0.032	Catanzaro	1.094	0.041
Cremona	0.863	0.033	Genova	1.095	0.029
Oristano	0.865	0.031	Ancona	1.103	0.026
Novara	0.866	0.031	Milano	1.104	0.026
Piacenza	0.868	0.039	Firenze	1.106	0.024
Vercelli	0.874	0.036	Pescara	1.117	0.030
Torino	0.879	0.030	Pisa	1.153	0.022
Teramo	0.883	0.030	Udine	1.156	0.030
Savona	0.888	0.030	La Spezia	1.159	0.032
Cosenza	0.890	0.036	Campobasso	1.161	0.032
Forlì-Cesena	0.900	0.029	Bologna	1.185	0.027
Trapani	0.905	0.026	Siena	1.208	0.022
Cuneo	0.908	0.027	Cagliari	1.340	0.036
Ravenna	0.914	0.031	Trento	1.368	0.031
Benevento	0.920	0.030	Trieste	1.382	0.033
Ascoli Piceno	0.925	0.030	Bolzano-Bozen	1.421	0.036
Verbano-Cusio-Ossola	0.927	0.030	L'Aquila	1.484	0.034
Matera	0.932	0.043	Valle d'Aosta/Vallée d'Aoste	1.545	0.039
Caltanissetta	0.936	0.034	Roma	1.775	0.027

Notes: The province shares are ratios of the average government spending in province *i* to the average national government spending. Sample standard deviations for each cross-sectional unit are computed to obtain the GDP volatility for province *i*.

TABLE A2 Local projections of the loans to government spending shock.

	Total loans			
Horizon	**0**	**1**	**2**	**3**
Coeff.	0.0097#	0.0195***	0.0186***	0.0049#
Std. errors	(0.0055)	(0.0015)	(0.0024)	(0.0026)
	Loans to NFCs			
Horizon	**0**	**1**	**2**	**3**
Coeff.	0.0026	0.0126***	0.0131***	−0.0051
Std. errors	(0.0042)	(0.0027)	(0.0035)	(0.0033)
	Loans to small businesses			
Horizon	**0**	**1**	**2**	**3**
Coeff.	−0.0009	0.0011#	0.0021***	0.0011***
Std. errors	(0.0009)	(0.0006)	(0.0005)	(0.0002)
	Loans to consumer households			
Horizon	**0**	**1**	**2**	**3**
Coeff.	0.0071***	0.0069***	0.0055***	0.0101***
Std. errors	(0.0020)	(0.0020)	(0.0019)	(0.0010)

Notes: Driscoll and Kraay (1998) robust standard errors in brackets.
***, **, *, # indicate statistical significance at 0.1, 1, 5 and 10% levels.
See Equation (7) and Figure 6.

TABLE A3 Local projections of the loans to government spending shock: distinction by macro-geographical area.

	Total loans			
Horizon	**0**	**1**	**2**	**3**
Centre–North	0.0147*	0.0242***	0.0261***	0.0163**
South	−0.0349*	−0.0201	−0.0465*	−0.0960***
	Loans to NFCs			
Horizon	**0**	**1**	**2**	**3**
Centre–North	0.0055	0.0151***	0.0185***	0.0030
South	−0.0177	−0.0027	−0.0273	−0.0786***
	Loans to small businesses			
Horizon	**0**	**1**	**2**	**3**
Centre–North	−0.0006	0.0015	0.0029***	0.0022***
South	−0.0040#	−0.0032	−0.0069*	−0.0082#
	Loans to consumer households			
Horizon	**0**	**1**	**2**	**3**
Centre–North	0.0093***	0.0090***	0.0076**	0.0133***
South	−0.0172***	−0.0174***	−0.0192***	−0.0174*

Notes: ***, **, *, # indicate statistical significance at 0.1, 1, 5 and 10% levels.
See Equation (7) and Figure 7.

TABLE A4 Local projections of the loans to government spending shock: by area and bank headquarters.

	Banks in the Centre–North and GDP			
Horizon	**0**	**1**	**2**	**3**
Centre–North	−0.0203	0.0114	−0.0262	−0.0236
South	−0.1273*	−0.0683	−0.2055***	−0.1082***
	Banks in the South			
Horizon	**0**	**1**	**2**	**3**
Centre–North	0.0049***	0.0020#	0.0040	0.0035**
South	0.0667*	0.0423	0.0962*	0.0665***

Notes: ***, **, *, # indicate statistical significance at 0.1, 1, 5 and 10% levels.
See equation (7) and Figure 8.

TABLE A5 Robustness checks for the first stage, using total GVA at constant prices instead of real GDP as a measure of the real economic activity.

	Government consumption (GVA)	
	Estimate	**Std. errors**
Bartik	0.6646***	0.0838
F-statistic	32.5439	
*** p-value < 0.001		
Estimator		**F-statistic**
White (1980)		27.1285
White (1984)		69.7509
Arellano (1987)		16.5714
Driscoll and Kraay (1998)		13.2699
Kleibergen-Paap		26.133

TABLE A6 Robustness checks for Equation (7), using total GVA at constant prices instead of real GDP as a measure of the real economic activity.

	Total loans			
Horizon	**0**	**1**	**2**	**3**
Coeff.	0.0108#	0.0216***	0.0203***	0.0057*
Std. errors	(0.0061)	(0.0016)	(0.0026)	(0.0029)
	Loans to NFCs			
Horizon	**0**	**1**	**2**	**3**
Coeff.	0.0030	0.0140***	0.0142***	−0.0054
Std. errors	(0.0047)	(0.0030)	(0.0038)	(0.0037)
	Loans to small businesses			
Horizon	**0**	**1**	**2**	**3**
Coeff.	−0.0010	0.0012#	0.0023***	0.0012***
Std. errors	(0.0010)	(0.0007)	(0.0006)	(0.0002)
	Loans to consumer households			
Horizon	**0**	**1**	**2**	**3**
Coeff.	0.0078***	0.0075***	0.0061**	0.0111***
Std. errors	(0.0021)	(0.0022)	(0.0021)	(0.0011)

Notes: Driscoll and Kraay (1998) robust standard errors in brackets.
***, **, *, # indicate statistical significance at 0.1, 1, 5 and 10% levels.

TABLE A7 Robustness checks for Equation (7) for the two sub-samples, Centre–North and South.

	Total loans			
Horizon	**0**	**1**	**2**	**3**
Centre–North	0.0162*	0.0267***	0.0285***	0.0183**
South	−0.0378*	−0.0232	−0.0531*	−0.1081***
	Loans to NFCs			
Horizon	**0**	**1**	**2**	**3**
Centre–North	0.0060	0.0168***	0.0201***	0.0038
South	−0.0187	−0.0035	−0.0313	−0.0889***
	Loans to small businesses			
Horizon	**0**	**1**	**2**	**3**
Centre–North	−0.0007	0.0016	0.0032***	0.0024***
South	−0.0044#	−0.0037	−0.0077*	−0.0094#
	Loans to consumer households			
Horizon	**0**	**1**	**2**	**3**
Centre–North	0.0102***	0.0099***	0.0084**	0.0145***
South	−0.0191***	−0.0197***	−0.0218***	−0.0192*

Note: ***, **, *, # indicate statistical significance at 0.1, 1, 5 and 10% levels.

TABLE A8 Robustness checks for Equation (7) by area and bank headquarters.

	Banks in the Centre–North and GDP			
Horizon	**0**	**1**	**2**	**3**
Centre–North	−0.0223	0.0130	−0.0301	−0.0255
South	−0.1539**	−0.0827	−0.2270***	−0.1350***
	Banks in the South			
Horizon	**0**	**1**	**2**	**3**
Centre–North	0.0055***	0.0023*	0.0044	0.0040**
South	0.0882***	0.0522#	0.1012*	0.0869***

Note: ***, **, *, # indicate statistical significance at 0.1, 1, 5 and 10% levels.

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Resumen. Este estudio examina los efectos de las perturbaciones del gasto público en el mercado de crédito italiano, para lo cual utiliza datos NUTS 3 con un periodo de muestra de 2011 a 2018. La metodología empírica se basa en una proyección local IV y la identificación de una perturbación del gasto público se logra construyendo un instrumento Bartik. La evidencia empírica muestra un efecto positivo leve del aumento del 1% del gasto público en relación con el PIB sobre el crecimiento del volumen de préstamos en relación con el PIB. Sin embargo, los resultados empíricos muestran que el gasto público no ayuda a mejorar ni el “sesgo de tamaño”, es decir, las restricciones financieras a las que se enfrentan las pequeñas empresas en relación con las más grandes, ni el “sesgo doméstico” en los préstamos relacionados con el proceso de consolidación bancaria en Italia.

抄録: 本稿では、政府支出ショックがイタリアの信用取引市場に及ぼす影響を、2011~2018年の期間のNUTS 3のデータを用いて検討した。Local Projection IVに基づいた実証的方法を用いて、バーティク操作変数を作成し公共支出ショックを特定する。実証的に得られたエビデンスから、GDPに対して政府支出が1%増加すると、GDPに対する融資額の比率が増加するという、わずかなプラス効果が示される。しかし、実証的に得られた知見によると、政府支出は、大企業よりも小企業が直面する財務的制約である「規模バイアス」や、イタリアの銀行統合プロセスに関連する融資における「ホームバイアス」 のいずれの改善にも役立たないことが示される。