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Essays on fiscal policy shocks in Italy: empirical analysis using regional data

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Introduction

In the last twenty years there have been remarkable episodes of financial turmoil: the global financial crisis of 2008 and the sovereign debt crisis in 2011 which had significant consequences on Italy and spread differently across regions. All of these crises posed difficult challenges for policymakers, and in particular triggered a discussion and rethinking of the use of fiscal policy, to be used alongside monetary policy, to deal with the consequences of these crises. In this context, it is worth recalling a famous statement by former ECB President Mario Draghi: "it's now high time I think for the fiscal policy to take charge". The intensive use of monetary policy tools to deal with these crises, which have seen interest rates reach close to zero and even negative values, requires an alternative tool to further stimulate the economy and foster recovery from crises. Therefore, the interest of policymakers and academics on the effects of fiscal policy has gained attention (Gabriel et al., 2023). Moreover, this resulted in the implementation of several fiscal stimulus packages after the financial crisis and the Covid-19 pandemic. In particular, the latter episode stimulated the willingness of Eurozone countries to engage, for the first time, in a new EU budget allocation that not only takes funds from national contributions, but also from international financial markets (Canova and Pappa, 2021).

The focus of this thesis is on the empirical analysis of local fiscal policy effects in Italy, given the historical disparities, in terms of economic development, between Centre North and Southern regions. Given a relatively short time series data span (involving at most the last two decades) the focus is not on long term growth but on the dynamic of output at business cycles horizon and on the impact on credit market conditions and housing market as main drivers of output fluctuations. I analyze both the average impact of local fiscal policy and, in order to assess whether there is evidence of redistributive effects during period of financial turmoil, I also examine the different effects of fiscal policy on Centre-North and South macro-regions. Moreover, the focus is on the unanticipated exogenous fiscal policy effects through identification of structural shocks hitting mainly government spending using panel data at the NUTS-2 (which, according to the official classification of the European Commission, corresponds to 19 Italian regions plus the two autonomous provinces of Bolzano/Bozen and Trento) or NUTS-3 level (involving 106 Italian provinces).

I focus on two main categories of public spending, namely nationally financed government spending and spending channeled through the European Structural Funds, particularly investment spending financed by the European Regional Development Fund (ERDF).

Recently, the Italian Institute of Statistics (ISTAT) and the European Commission have provided various data at the NUTS-2 and NUTS-3 level. Therefore, I exploit these sources to obtain data on different proxies of output and government spending. In particular, the ARDECO database of the European Commission provides data on GDP and GVA, for 6 NACE sectors, at the NUTS-2 and NUTS-3 level. In addition, ISTAT provides government expenditure, GDP and GVA at the NUTS-2 level. The EU historical payments database by NUTS-2 regions was used to retrieve data on EU Structural Funds allocations. Furthermore, more detailed data on the public accounts of the Italian regions are available from the Government Agency for Territorial Cohesion ("Conti Pubblici Territoriali"), from which data on government spending and revenues can be obtained. Another important source, which provides detailed data on the credit market at different territorial levels, is the Bank of Italy's BDS ("Base Dati Statistica") database. All these sources are publicly available. However, I also use confidential data on house prices, obtained from the real estate observatory market of the Italian Revenue Agency ("Osservatorio del Mercato Immobiliare" or "OMI").

In the first chapter, I focus on the effects of public spending shocks on the credit market at the local level in Italy. In a study involving US counties, Auerbach et al. (2020) show that the expansion of government spending, in contrast to standard macroeconomic models (which predict a rise in interest rates), can relax credit market conditions. This can happen because local government spending can be interpreted as an injection of liquidity into local economies, which can have two effects. First, if the local credit markets are segmented, this injection can lower interest rates in that county. Moreover, regardless of whether local credit markets are segmented or integrated across regions, the positive effects that government spending can have on that region's economy can improve its economic conditions and, for example, lower the probability of a local recession, thus reducing the credit risk in that region and easing borrowers' access to the credit market. Thus, although standard theory has emphasized the "counterproductive" role of fiscal policy to stimulate credit markets, recent evidence points to its positive effects (an extensive review of the effects of fiscal policy on credit market conditions and further empirical evidence can be found in Murphy and Walsh, 2022). This leads to the analysis conducted in the first chapter, in which I exploit data on 106 Italian NUTS-3 provinces over the period 2011-2018, to estimate the effects of public consumption shocks on bank lending to the private sector in the Italian provinces. The short sample period (in terms of time series dimension) is a period characterised by the post-global financial crisis (GFC), the sovereign debt crisis and the monetary policy constrained at the Zero Lower Bound (ZLB). The focus of the analysis is not only on the average effects of local fiscal policy across provinces. I also examine heterogeneous local fiscal policy effects. First, I test whether the "size bias" related to the financial constraints of small firms is mitigated by increases in public spending, observing how credit granted to small firms reacts to public spending shocks, compared to larger firms. Moreover, I test the "home bias". In particular, given the process of banking consolidation that has given a central role in the banking system to the larger banks located in the North of Italy, I assess whether there is evidence of an increase in the so-called functional distance, exacerbating the financial constraints faced by small firms located far from the banks' headquarters. More specifically, I study how the effects of public spending shocks change depending on the area in which borrowers and lenders are located. To do so, I use the Local Projections approach proposed by Jordà (2005), to estimate the dynamic responses of loans to government spending shocks, which are identified using an Instrumental Variable known as "shift-share" or Bartik-type instrument (Bartik, 1991), which essentially consists of decomposing the local public expenditure in a constant provincial factor, which is represented by the average share of provincial government spending in the national expenditure, and a time-varying factor, which is given by changes in national government expenditure. This IV has been widely used in the empirical literature to identify government spending shocks in a panel of regions or industries (see, e.g.: Nekarda and Ramey, 2011; Nakamura and Steinsson, 2014; Auerbach et al., 2020; Gabriel et al., 2023). The empirical findings suggests that government spending stimulates the local credit markets in Italy, but the effects are higher for the whole Non-Financial business sector relative to smaller firms and households. I also document a reduction in the risk of borrowers by showing that the government spending shock lowers the bad loans rate. Furthermore, the positive effects prevail in the more developed area of the Centre-North. Finally, fiscal policy does not help to reduce the "home bias" related to the banking consolidation process, highlighted above, since the positive effects of government spending shocks on the credit market of the "Mezzogiorno" are only found when loans are provided by local banks.

In the second chapter I focus on a different category of public expenditure, which is the investment expenditure implemented under the program of the European Regional Development Fund (ERDF). This fund invests on Information and Communication Technologies, Research and Development and provides funds to Small and Medium-Sized Enterprises. The focus is, first, on the analysis of the effects of the ERDF transfers on the economy of the Italian regions, by estimating output multipliers. Moreover, I examine (in a second stage) whether government spending (funded through ERDF funds) has played a countercyclical role in ameliorating the financial fragility of the private sector during the most recent period related to global financial crisis and sovereign debt crisis. The panel dataset used includes the ERDF transfers and three proxies of real output, namely, real GDP, real GVA and real GVA of the private sector, in the 21 Italian NUTS-2 regions over the period 1988-2018. The identification method of fiscal spending shocks follows Brueckner et al. (2023) and Canova and Pappa (2021), by retrieving a proxy for the shock hitting the ERDF transfers as the residual of a reaction function, where the ERDF transfers are regressed on current and past output and past ERDF transfers. Current output is instrumented in this regression with the GDP prevailing in the macro area (NUTS-1) to which each region belongs, excluding the GDP of the region. However, unlike Brueckner et al. (2023) and Canova and Pappa (2021), this residual is used in the proxy-SVAR (see Stock and Watson, 2012; Mertens and Ravn, 2013; Gertler and Karadi, 2015) which is applied to panel data. The empirical evidence reveals that ERDF transfers have significant effects on the Italian regions, with cumulative multipliers over three years ranging from 1.17 for private sector GVA, to 2 for total GVA and 2.28 for GDP. Then, the identified fiscal spending shock is used to assess its effects on the financial fragility of the business sector, proxied by the Non-Performing Loans to potential output ratio, in the Italian provinces over the financial and sovereign debt crisis period (2009-2018). For this purpose, I use the Local Projections approach (Jordà, 2005), to estimate the dynamic responses of the NPL-to-output ratio to a regional ERDF shock. As a result, I show that the ERDF shocks significantly lower the NPL-to-output ratio, especially for the construction and manufacturing sectors, where much of the ERDF investments are concentrated, especially in the manufacturing sector (see Canova and Pappa, 2021). Further investigation, through non linear local projections (via Smooth Transition, with a local credit diffusion index provided by the Bank of Italy used as transition variable), shows that the beneficial effects on the financial fragility of the business sector due to ERDF fiscal spending shocks prevail during a credit supply easing regime. Finally, the non linear local projections (via Smooth Transition using the Regional Competitiveness Index, provided by the European Commission, as transition variable) shows that policy makers, beyond the use of government guarantees granted to banks, should take into account the beneficial effect of ERDF fiscal spending improving the financial conditions of the business sector especially in regions with a low degree of competitiveness.

In the third chapter I exploit the methodology developed in chapter two to estimate the output multipliers of nationally financed government expenditure in the 21 Italian NUTS-2 regions. Indeed, the third chapter uses a panel dataset containing 21 Italian regions over the period 1995-2019, extending the reaction function used in the second chapter with public expenditure forecasts to obtain a proxy for the regional government spending shock. Government spending is regressed on current and past output, past government spending and on the forecast of government spending. This allows to obtain a series of changes in regional government spending which are unrelated to current and past economic conditions and purged from the part of the spending which is anticipated. As in the second chapter, the output in this regression is instrumented by a variable constructed as the interaction between changes in the international oil prices and the average regional share of manufacturing value added in the total value added. The application of the proxy-SVAR delivers cumulative government spending multipliers equal to 1.26. Because of the different economic and structural conditions characterizing the North and the "Mezzogiorno", I extend the model, using a dummy taking 1 if the region belongs to the "Mezzogiorno" and 0 otherwise, to check whether the multipliers are different in these two macro-areas. I estimate a dummy-augmented VAR model to obtain distinct parameters for the two macro-areas. The results point at higher effectiveness of fiscal policy in the more developed area of the Centre-North. The six years cumulative government spending multipliers is found to be 1.65 in the Centre-North, against a value of 0.91 in the "Mezzogiorno". However, the difference in the multipliers between the two macro-areas is statistically significant in the short-run only.

Finally, in the last chapter, I turn my focus on the study of how housing markets, together with household credit market conditions, respond to fiscal policy shocks, both on the expenditure and revenue side. In particular, after having documented the positive effects of public spending on the economy and credit markets and the reduction of borrowers' financial fragility and credit risk in the previous chapters, I analyse how house prices react to public spending and tax shocks at the regional level for Italy, and I also study the dynamic response of credit market conditions for households (different categories of credit granted to households and interest rates applied by banks on loans to households) after tax revenues shocks. House prices are modelled as endogenous variables, since fiscal policy effects on the housing market depends not only on the credit channel, but also on the role played by housing property as collateral to obtain new loans, reducing the financial constraints that households face, thus helping them to access the credit market (Khan and Reza, 2017). The empirical literature on the effects of fiscal policy on house prices has mainly concentrated on studies involving country-level data, and without any focus on credit market conditions both in terms of types of households loans and interest rates applied to households loans (see Afonso and Sousa, 2011-2012; Agnello and Sousa, 2011; Khan and Reza, 2013-2017; Gupta et al., 2014; Aye et al., 2014; Ruiz and Vargas-Silva, 2016). I construct a panel dataset at the NUTS-2 level for Italy over the period 2004-2019. It includes data on government expenditure, government revenues, output, house prices, bank loans granted to households for different purposes, such as, mortgages, loans for the purchase of durable goods, and consumer credit, and long-term interest rates charged by banks on households loans. Government spending and tax shocks are then identified through sign restrictions (Uhlig, 2005), following Canova and Pappa (2007) and Pappa (2009). I employ a Bayesian VAR analysis, exploiting the methodology proposed by Banbura et al. (2007), incorporating prior information by means of so-called artificial data, which are created in a proper way in order to impose a Normal prior on the panel VAR coefficients and an Inverse-Wishart prior on the covariance matrix. The baseline results suggest that an expansionary fiscal policy stimulates the housing markets in the Italian regions, increasing house prices. At the same time, fiscal expansion increases income and eases credit market conditions, lowering interest rates for households. Furthermore, the data also show that the volume of loans increases after an expansionary fiscal shock. To check how the different level of economic development and banking sector development of the Italian regions shape the way in which house prices and households credit market conditions react to fiscal policy shocks, I estimate the model in sub-samples of our dataset. Moreover, I examine whether the fiscal policy effects could have changed after the Global Financial Crisis. Indeed, this seems to be the case, since the estimation of the model on the two sub-samples involving the period before and after the Global Financial Crisis shows that fiscal policy became more effective in stimulating the housing market and relaxing the credit market conditions after the Global Financial Crisis, therefore acting in a counter-cyclical way for both the housing market and the credit market, that were hit by bubbles and severe crises during that period. However, further empirical analysis does not suggest a redistributive role of fiscal policy beneficial to Southern regions. More specifically, first, I divide the sample according to the level of economic development and I find that the positive effects documented in the baseline analysis prevail in the more developed regions, where I observe higher effects of fiscal policy shocks on the variables of interest. Then, I split the regions depending on the level of banking sector development, and the analysis shows that the fiscal policy effects on house prices and credit market conditions are stronger in the group of regions with a higher level of banking sector development, where the credit channel may work better.

Overall, the empirical analysis of this thesis suggests, controlling for unobserved heterogeneity across Italian regions/provinces, a positive average effects of fiscal policy on output and a reduction in the credit risk and financial fragility of borrowers (in line with Auerbach et al. 2020), and thus an easing of credit market conditions, with further positive effects on house prices. However, there is considerable heterogeneity of fiscal policy effects across two macro-areas of the country, i.e. the North and the South, and this calls for a targeted policy to reduce the gap between these two very different areas. Although this is outside the scope of this analysis, I show that this gap has important consequences, and this topic can thus be a stimulus for future research. Moreover, another feature of the analysis is that it is based on a homogeneous panel approach used throughout the thesis, leaving only fixed effects to control for heterogeneity. First, as emphasised in the literature focusing on the estimation of the so-called *"geographic cross-sectional fiscal spending multipliers"* (see, among others, Nakamura and Steinsson, 2014 and Chodorow-Reich, 2019), the use of time fixed effects allows to control for common shocks, such as monetary policy, which is one of the main confounding factors in this type of analysis. As pointed out by Nakamura and Steinsson (2014), this is a great advantage in estimating the multiplier in this context. Secondly, for most of the datasets I constructed, the panel time series is very short, which limits the use of other estimation strategies such as Pesaran and Smith's (1995) popular mean group estimator. I am aware of the existence of partial pooling methods that can potentially solve the sample size problem and I am still exploring this area, which I intend to apply in future developments of this thesis. I also recognise that there are important issues related to spillover effects between regions. This type of analysis deserves particular attention especially when the sample size T, regarding the time series dimension, is short. Canova and Pappa (2007) argue that the shortness of the dataset may prevent the use of richer models to study the transmission of shocks between units. This sometimes forces researchers to make choices on the restrictions imposed on the coefficients of the panel VAR model. Therefore, all these points need to be further developed in future research.

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Chapter 1

1 Government spending and credit market: evidence from Italian (NUTS-3) provinces

1.1 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 rate, 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 at 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. 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 to the bank assets and lower devaluation of the 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 can 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 Italian local credit markets. Italy is a relevant case study in this context because of the differences in structural and economic conditions between provinces located in different areas of the country. The gap between the most developed area in the Centre-North and some areas in the South, usually referred to as "Mezzogiorno", is remarkable. 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). This aspect is also relevant in our case, because small enterprises generate a considerable share of the overall value added of the Italian non-financial sector and rely more on external financing through the banking market, a characteristic that Italy shares with other Eurozone countries, where the financial system is bank-based. Third, 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. This further makes Italy an interesting case study, as the closure of banks in the South and the consequent polarisation of the banking system in the North can exacerbate the financial constraints faced by households and small businesses in Italian provinces located in the South.

The empirical analysis is based on a local projection 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 variation in government spending is achieved by constructing a Bartik (1991) type instrument which, according to Auerbach et al. (2020), allows 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 projections (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,

[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.

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 chapter is organized as follows. Section 1.2 provides a literature review on the empirical studies of the impact of government spending on credit markets. Section 1.3 describes the data, the empirical methodology, including the identification of government spending shocks, and the empirical evidence. Section 1.4 describes the robustness analysis and Section 1.5 gives some concluding remarks.

1.2 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 (Murphy and Walsh, 2022; Devereaux et al., 1996).

However, 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 World War I, 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 either 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 episode. 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

VAR model, involving data for the U.S. 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 U.S. 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'Alessandro et al. (2019) use a 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 a structural VAR 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 (Bernoth et al., 2003; Codogno et al., 2003; Afonso and Strauch, 2003; Burriel et al., 2009). 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.

1.3 Empirical analysis

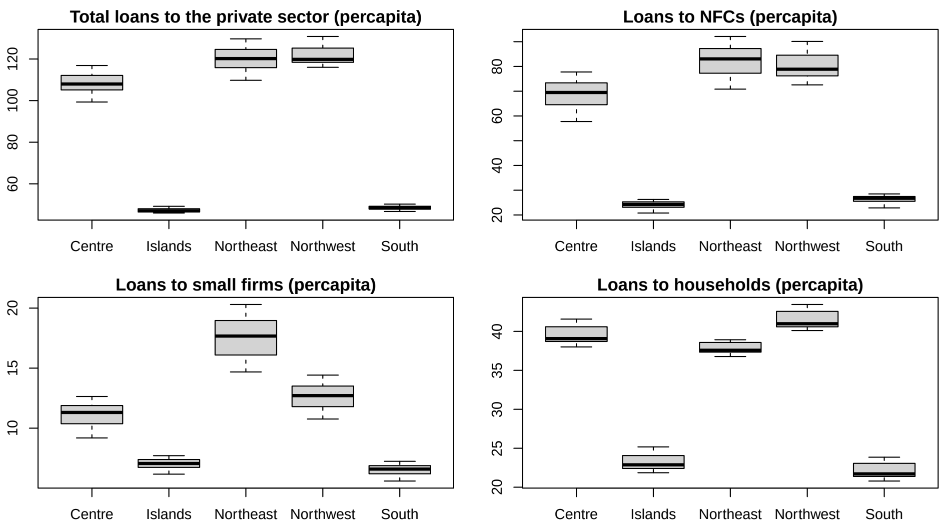
1.3.1 Data

We merge data from the Bank of Italy 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 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 1.1 and Figure 1.1 provide descriptive statistics and the boxplot of the distribution of loans across NUTS-1 regions, respectively, that give 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, NFCs and households allocated to the former.

Table 1.1: 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

Figure 1.1: Boxplots of loan distribution across NUTS-1 regions (euro per capita).



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

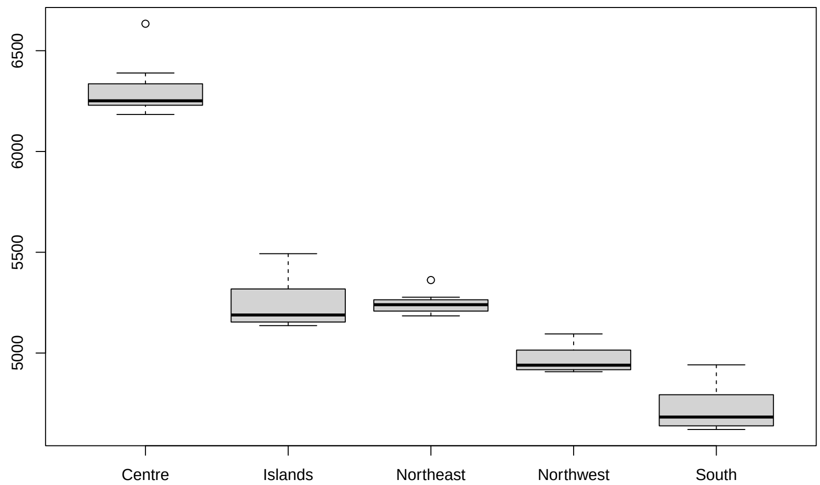
[2] We are aware of the short time series dimension of the panel limiting the empirical analysis to the period during and after the sovereign debt crisis. However, in order to investigate the risk premium channel as the main transmission mechanism from government spending to credit (as suggested by Auerbach 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.

data for government spending we follow the suggestions of Gabriel et al. (2023) and Brueckner et al. (2023) and we use the Gross Value Added 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 Gross Value Added 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 defense; (ii) Education; (iii) Human health and social work; (iv) Arts, entertainment, and recreation; (v) Other service activities; (vi) Activities of households and extra-territorial organizations and bodies. As pointed by Gabriel et al. (2023) and Brueckner et al. (2023), the Gross Value Added 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 1.2: 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 1.2: Boxplot of the distribution of the government spending proxy in NUTS-1 regions (millions of euro per capita, real terms).



[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. (2023) 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.

Table 1.2 and Figure 1.2 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 1.3 shows the two series in log form, and Figure 1.4 shows the two series of the first difference of these two variables.

Figure 1.3: Plot of the time series of the Final Consumption Expenditure of the General Government and the GVA of the Non-market Sector in logarithm.

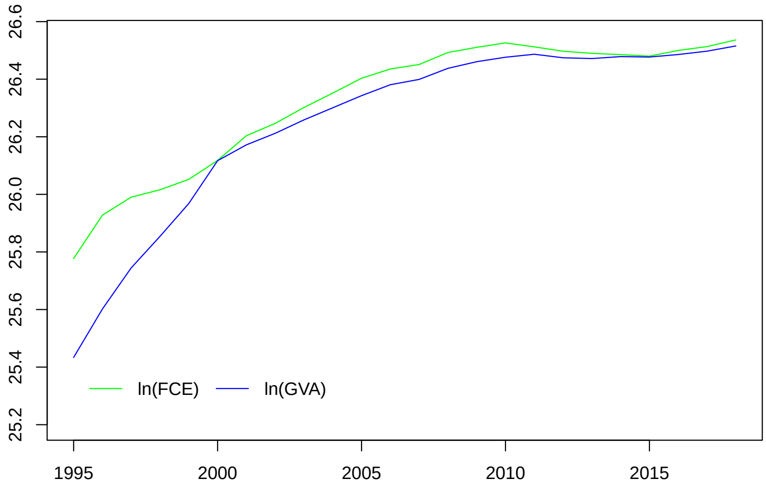
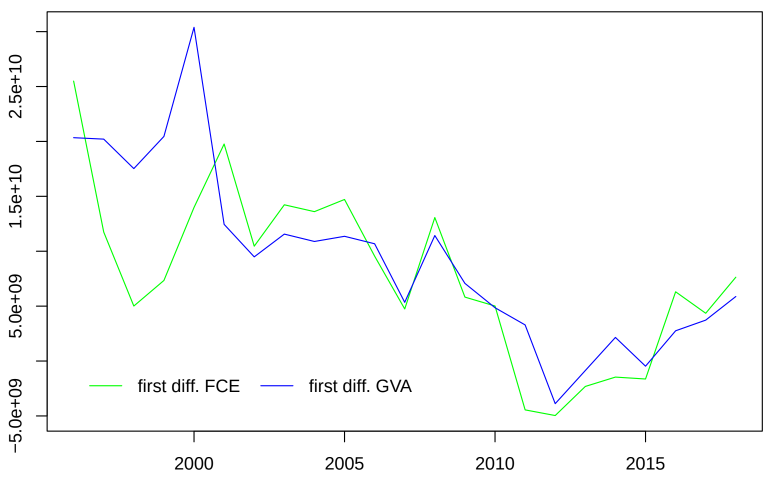


Figure 1.4: Plot of the time series of the first difference of the Final Consumption Expenditure of the General Government and the GVA of the Non-Market Sector.



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. Table 1.3 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 1.4 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 heteroskedasticity and serial correlation. Thus, the GVA of the non-market sector seems to be a good proxy for government spending at the provincial level.

Table 1.3: Pearson Correlation between Government Spending proxies from AMECO and ARDECO.

$Corr(ln(FCE_t), ln(GVA_t))$	$Corr(\Delta FCE_t, \Delta GVA_t)$
0.9838***	0.7115***

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

Table 1.4: Regression analysis for Government Spending proxies from AMECO and ARDECO.

	$ln(FCE_t)$
$ln(GVA_t)$	1.0028*** (0.0014)
	ΔFCE_t
ΔGVA_t	0.7557*** (0.1547)

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

Finally, it is worth noticing 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 sub-section 1.3.3).

1.3.2 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 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):

$$\frac{L_{i,t+h} - L_{i,t+h-1}}{Y_{i,t-1}} = \alpha_{i,h} + \gamma_{t,h} + \beta_h \frac{G_{i,t} - G_{i,t-1}}{Y_{i,t-1}} + \delta_h X_{i,t} + \epsilon_{i,t+h} \tag{1.1}$$

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 heteroskedasticity and autocorrelation using an HAC estimator to calculate the standard errors. In line with Furceri et al. (2021), Gabriel et al. (2023) and Auerbach et al. (2020), we use the Driscoll and Kraay (1998) estimator, which not only controls for heteroskedasticity 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., 2023; Nakamura and 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. Since the short data span (due to availability of the loans to small firms dataset) involves only eight 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.

1.3.3 Identification strategy

We acknowledge the endogeneity of government spending in equation (1.1). Indeed, apart from the omitted variable issue, the endogeneity may results from reverse causality between government spending and loans. Both central and local governments may take into account credit market developments and decide the amount of spending accordingly, either because they are concerned about credit market conditions or because of the effects that shocks to the credit market may have on the economic activity.

Given the small T large N feature of the panel dataset used, we cannot rely on the identification schemes implemented in a Structural Vector Autoregression, 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. (2023) (who focus on European government spending data) we construct an instrumental variable for government spending by, first, constructing the time invariant share[4]:

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

$$s_i = \frac{\bar{G}_i}{\bar{G}_{ITA}} \tag{1.2}$$

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, the 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:

$$g_t = \frac{G_{ITA,t} - G_{ITA,t-1}}{Y_{ITA,t-1}} \tag{1.3}$$

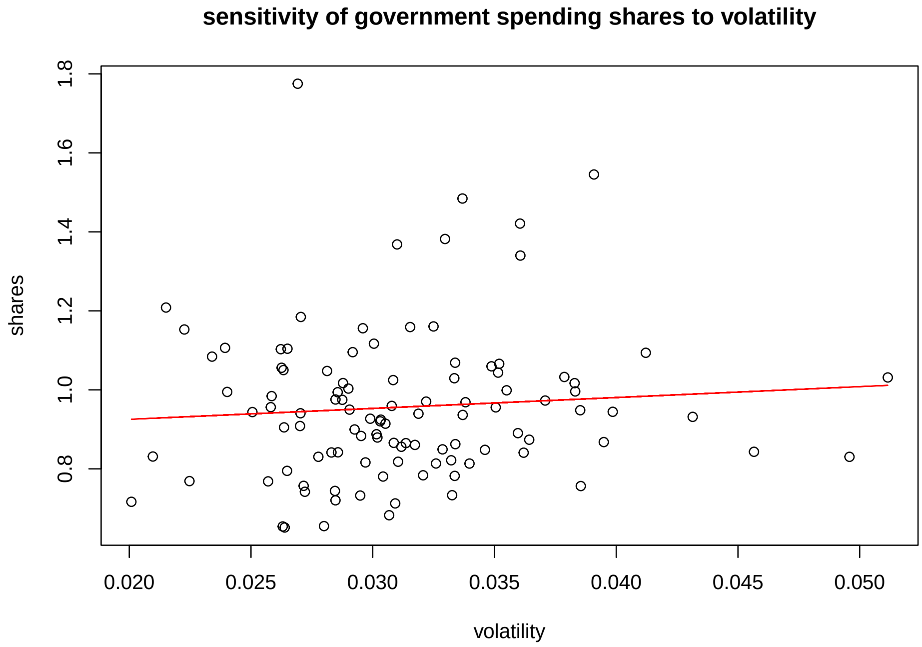
Finally, the instrument is constructed by computing the Kronecker product between the vector containing the shares and the vector containing changes in national government spending, which corresponds to the multiplication of the share of each province by the changes in national expenditure in each year:

$$\begin{gathered} \mathbf{Bartik} = \mathbf{s} \otimes \mathbf{g} \\ Bartik_{i,t} = s_i \times g_t \end{gathered} \tag{1.4}$$

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, i.e., the violation of the identifying assumption, could come from the vector containing the shares. However, following Gabriel et al. (2023) and also Nakamura and Steinsson (2014), we verify if our IV suffers from the endogeneity issues just described.

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 and Steinsson, 2014). We note that the standard deviation of each local unit's real GDP growth does not change substantially across units and, in particular, it shows similar values for local units with shares above and below the median (see Table A1.1 in Appendix A1). 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 1.5).

Figure 1.5: Scatterplot of shares and volatility of real GDP growth.



Note: the red line is the estimated regression line

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. We can see that there is no obvious relationship between the shares and volatility. The results of the regression analysis are in Table 1.5.

Table 1.5: Test on the instrument a).

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

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

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. (2023), 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, i.e., regions receiving a larger volume of public spending, become poorer than other regions. Table 1.6 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 section 1.3.1, 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.

Table 1.6: Test on the instrument b).

	Bartik Instrument	
	Pooled OLS	Within FE
Intercept	0.003***	
	(0.0001)	
RSBC	0.003	0.002
	(0.003)	(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.
Note: The results in the table show the sensitivity of the Bartik instrument to the measure of the relative stance of the business cycle (RSBC).
*** indicate statistical significance at 0.1% level.

Thus, we conclude that our identification assumption, relying on the exogeneity of the instruments, is valid[5].

Tests for the relevance of the instrument So far, we have discussed only the exogeneity assumption and tried to give some evidence in favor of it. However, another important condition of the instrumental variable that must be satisfied is the relevance, meaning that the instrument must be highly correlated with the endogenous variable. More specifically, the relevance condition can be written into the formula as follows:

$$E\left[\left(\frac{G_{i,t}-G_{i,t-1}}{Y_{i,t-1}}\right)\times Bartik_{i,t}|\alpha_i,\gamma_t,X_{i,t}\right]\neq 0 \tag{1.5}$$

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

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 heteroskedasticity but without serial correlation; (ii) White's (1984) correction but assuming constant variance within groups; (iii) Arellano's (1987) estimator to control for both heteroskedasticity and serial

[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.

correlation; (iv) Driscoll and Kraay's (1998) estimator to control for heteroskedasticity, 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. (2023) 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 1.7 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.

Table 1.7: 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)		27.05
White (1984)		79.14
Arellano (1987)		15.60
Driscoll and Kraay (1998)		13.63
Kleibergen-Paap		26.06

1.3.4 Local Projections IV and size bias

In the first stage of the IV Local Projections 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:

$$\frac{G_{i,t} - G_{i,t-1}}{Y_{i,t-1}} = \alpha_i + \gamma_t + \beta Bartik_{i,t} + \delta X_{i,t} + \epsilon_{i,t} \tag{1.6}$$

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 changes 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.

[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; (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.

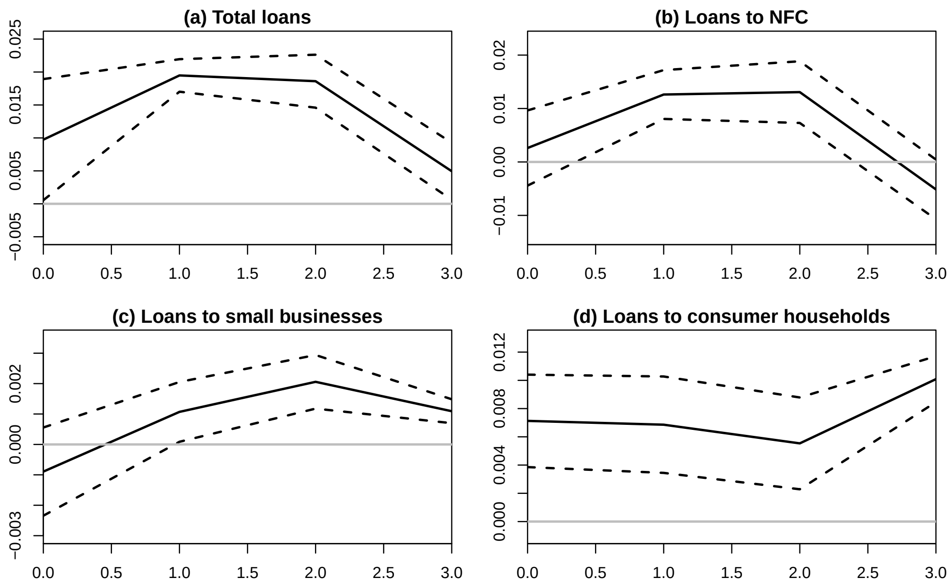
In the second stage of the Local Projections IV, we collect the fitted values from the first stage regression, that is $\frac{G_{i,t} - \hat{G}_{i,t-1}}{Y_{i,t-1}}$, only for the sub-sample 2011-2018 to match the availability of credit market data, and we estimate Local Projections regression equation:

$$\frac{L_{i,t+h} - L_{i,t+h-1}}{Y_{i,t-1}} = \alpha_{i,h} + \gamma_{t,h} + \beta_h \frac{G_{i,t} - \hat{G}_{i,t-1}}{Y_{i,t-1}} + \delta_h X_{i,t} + \epsilon_{i,t+h} \tag{1.7}$$

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

Figure 1.6 shows the Impulse Response Functions estimated by equation (1.7) (Table A1.2 in Appendix A1 shows the results).

Figure 1.6: Impulse Response Functions and 95% confidence interval bands.



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 one percent 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 households and on the small business sector is half and a fourth 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

[7] Specifically, Table 1.8 shows the results of the estimation of a panel regression similar to eq. (1.7), 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.

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 1.1, 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 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.

Table 1.8: 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*	-0.74**	-0.12
	(0.25)	(0.25)	(0.20)

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

1.3.5 Local Projections IV and home bias

Home bias. While the empirical evidence of Gabriel et al. (2023) 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", i.e., the geographical and economic distance between the banks' headquarters, i.e., 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", i.e., information that cannot be retrieved only by analyzing 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

[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.

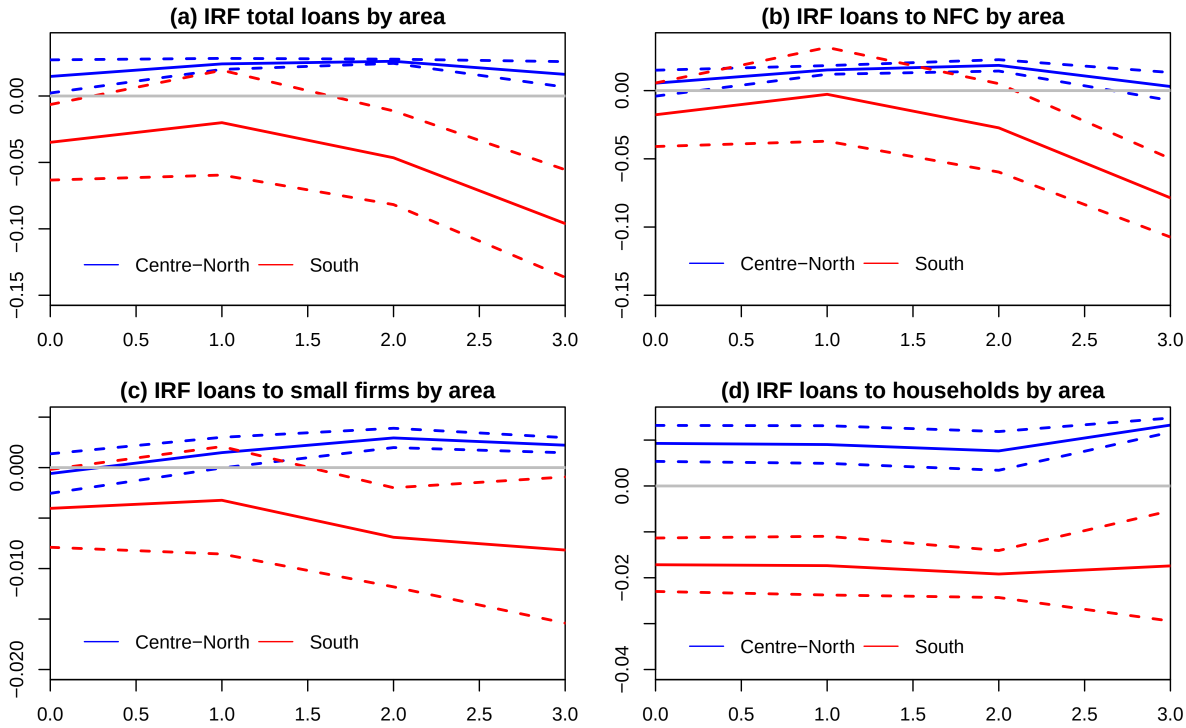
banks into local credit markets exacerbates the credit crunch in the post-Leman period[9]. Therefore, as explained in the Introduction, our third contribution to the literature focusing on the impact of government spending on credit growth, aims at evaluating whether the aforementioned home bias is reduced 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.

Empirical analysis of the geographical location of the borrowers. According to the classification of the Bank of Italy BDS Database, we consider the following territorial aggregation: (i) the northern regions are Piemonte, Valle d'Aosta, Liguria, Lombardia, Trentino-Alto Adige, Veneto, Friuli Venezia-Giulia, Emilia-Romagna; (ii) regions in the center are Toscana, Umbria, Marche, Lazio; (iii) the southern regions are Abruzzo, Molise, Campania, Puglia, Basilicata, Calabria, Sicilia, Sardegna. We divide our sample into two, one for the Centre-North and one for the Southern regions and estimate the Local Projections (1.7) to obtain the impulse response functions for each area. The IRFs are represented in Figure 1.7, whereas the results are in Table A1.3 in Appendix A1.

Figure 1.7(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 one percent 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 one, two and three 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 1.7, panel (b), (c) and (d)).

[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).

Figure 1.7: Impulse Response Functions by macro-geographical area and 95% confidence interval bands.



Empirical analysis of the geographical location of the banks. In the Bank of Italy 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 (1.7) 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 1.8 shows the Impulse Response Functions (Table A1.4 in Appendix A1 reports the results).

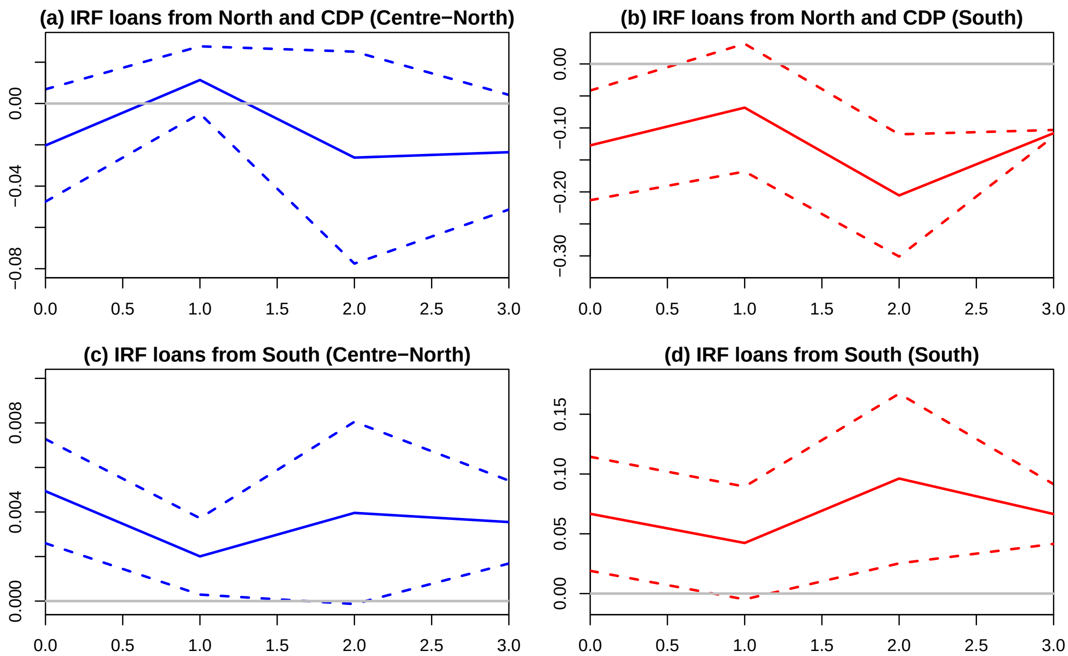
First, while the empirical findings in Figure 1.7 (see blue lines) show an increase in credit to private sector local borrowers in the Central-Northern provinces due to a public spending shock, the rightmost top panel in Figure 1.8 shows a negligible impact of this shock on the credit provided by banks 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 1.8, can be ascribed to a contraction 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).

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 one percent 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 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 one percent 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.

Figure 1.8: Impulse response functions by geographical macro-area and bank headquarters, and 95% confidence interval bands.



1.3.6 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 (1.7) and estimate the models in section 1.3.4 and 1.3.5. The results, shown in Appendix B1, are qualitatively similar, confirming the previous empirical findings.

[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 and Palley, 2002, 2017).

1.4 Conclusions

In this study we assess the effects of government spending on credit growth, employing the Local Projections 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 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", i.e., 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 SMEs 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 banking consolidation in Italy. Together, these results implies that government consumption, although it might be useful on aggregate to revitalize the credit market, it 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 suffers from violation of the exogeneity assumption once we move to a level of aggregation higher than NUTS-3. Finally, an important aspect that we do not consider in this study is the potential existence of spatial interactions between local credit markets (see Bellucci et al. (2013) showing that interrelationships between local credit markets may be important in Italy). We leave the investigation of the effects that a fiscal shock in one province may have on the credit markets of other provinces, due to cross-border relationships, for future research.

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Appendices to chapter 1

Appendix A1: Additional tables

Table A1.1: 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
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

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

Table A1.2: 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)

Note: See eq. (1.7) and Figure 1.6.
Driscoll and Kraay (1998) robust standard errors in brackets.
***, **, *, # indicate statistical significance at 0.1, 1, 5 and 10 percent levels.

Table A1.3: 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*

Note: See eq. (1.7) and Figure 1.7
***, **, *, # indicate statistical significance at 0.1, 1, 5 and 10 percent levels.

Table A1.4: 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***

Note: See eq. (1.7) and Figure 1.8
***, **, *, # indicate statistical significance at 0.1, 1, 5 and 10 percent levels.

Appendix B1: Robustness analysis

Table B1.1: 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)	27.13
White (1984)	69.75
Arellano (1987)	16.57
Driscoll and Kraay (1998)	13.27
Kleibergen-Paap	26.13

Table B1.2: Robustness checks for equation (1.7), 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)

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

Table B1.3: Robustness checks for equation (1.7) 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*

***, **, *, # indicate statistical significance at 0.1, 1, 5 and 10 percent levels.

Table B1.4: Robustness checks for equation (1.7) 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***

***, **, *, # indicate statistical significance at 0.1, 1, 5 and 10 percent levels.

Chapter 2

2 EU funds counter-cyclical effects: an analysis based on sub-regional data for Italy

2.1 Introduction

The focus of this study is the counter-cyclical role played by EU structural funds. Although, the primary target of EU cohesion policy is long-run growth convergence across EU regions, during the recent Covid-19 pandemic the emphasis has been on the reallocation of EU funds to a recovery from the crisis. EU funds finance public and private investments, which can produce direct and indirect demand-side effects in the short-term. This may justify an analysis of the short-term impact of EU structural funds on economic activity, through the estimation of Keynesian multipliers, and thus the exploration of countercyclical effects during economic and financial crises (see Di Caro and Fratesi, 2022; Psycharis et al., 2020; Neumark and Simpson, 2014)

Indeed, one stream of the literature on the EU cohesion funds counter-cyclical role has explored (across an extended time span) the effects of the EU funds over a short-medium term horizon, through the estimation of fiscal multipliers for a number of real economic activity proxies, mainly output (see Coelho, 2019, Canova and Pappa, 2021, and Destefanis and Di Giacinto, 2023 for the whole set of EU regions and Destefanis et al., 2022 for Italy). Another stream of the literature (Di Caro and Fratesi, 2022 and Di Pietro et al., 2021) has explored the counter-cyclical role of EU cohesion policy by assessing, through a reduced form model approach, whether EU structural funds can improve the resilience of output and labour market to crisis periods.

We contribute to the literature by exploring whether local public spending financed through the European Regional Development Fund (ERDF) has contributed to dampen the adverse consequences of crisis periods, such as the Great Recession and the sovereign debt crisis, on the corporate sector financial fragility across different territorial units in Italy. A further contribution is methodological since, in the first stage of the analysis, we identify, using an extended time span, the exogenous shock to the ERDF expenditure, in order to retrieve a proxy of discretionary policy intervention across different Italian NUTS-2 regions. In particular, we estimate a panel SVAR fitted to a dataset of 21 Italian NUTS-2 regions over the period 1988-2018 (annual observations). Contrary to existing panel VAR studies for Italy, using annual data available at NUTS-2 level, we do not rely on the Cholesky decomposition of the reduced form covariance matrix (see Deleidi et al., 2021 and Destefanis et al., 2022[12]), since a recursive identification scheme can be justified if the frequency of data is quarterly[13]. We argue that the

[12]Destefanis et al. (2022) also apply the Generalized IRFs approach as robustness check against the Cholesky identification.

[13]The study of Lucidi (2022) focuses on the estimation of multipliers for nationally financed fiscal spending, using a panel of Italian regions, employing a panel-BVAR estimation approach, where identification is achieved by imposing sign

zero exclusion restriction on the feedback effect of real economic activity on the policy variable is hard to reconcile with the main objective of EU structural funds, which is to stimulate regional economies, especially the poorest regions. Although the allocation of funds is decided before actual spending takes place, disbursements are determined by the demand for the investment projects submitted, which in turn means that actual spending is determined by the regions' business cycle conditions. Since we use a measure of the actual ERDF spending, we turn our focus on a different identifying scheme, based on the residual of a fiscal reaction function (we estimate a series of changes in the ERDF that are not related to GDP). Therefore, the additional methodological contribution to the literature is based on a Proxy SVAR methodology (extended to Panel data), which exploits additional moment conditions coming from the covariance equations between an instrument and the target shock (Stock and Watson, 2012; Mertens and Ravn, 2013; Gertler and Karadi, 2015). While the instrument used by Canova and Pappa (2021), to address the endogeneity bias of regional fiscal spending financed by EU funding, is given by using the innovations to EU structural funds spending orthogonal to macroeconomic conditions for the whole euro area and at country level, the instrument used in this study is the residual of a fiscal reaction retrieved by using data available at NUTS-2 level, and estimated in such a way that the innovations to the ERDF expenditure are orthogonal to past and contemporaneous region-specific macroeconomic conditions[14].

In the second stage of the analysis, we asses the effects of an exogenous ERDF spending increase on non-performing loans, over a short term horizon, to assess the role played by the ERDF in dampening the crises effects through their impact on the financial fragility of the private sector. This analysis, as in the first chapter, is motivated by Auerbach et al. (2020), who study the effects of fiscal policy on credit market conditions in the United States. The authors consider an increase in local government spending as an injection of liquidity into the local economies and argue that a transmission channel of government spending to credit market may work through the effect on the credit risk. Indeed, a spending expansion can lower the probability of a local economy to switch into a local recession, thus reducing the local credit risk. As discussed above, first, the demand-side effects of EU structural funds spending may also produce effects similar to those argued by Auerbach et al. (2020). In particular, EU funds, providing support for private investment and public investment, may directly and indirectly influence the credit market, through a reduction in the risk premium. (see Di Caro and Fratesi, 2022; Monfort et al., 2021). Secondly, the positive effects on the economy produced by ERDF investments can further strengthen these effects. In fact, the economic conditions of businesses can benefit from ERDF spending, with consequent positive effects on their balance sheets, thus reducing the financial fragility of the sector. Third, a part of the ERDF programme is dedicated to investments supporting small and medium-sized enterprises (SMEs), thus providing a direct channel to influence their economic conditions and indirectly reduce their financial fragility. This is important in the Italian context, where SMEs play a key role within the productive sector. This view is alternative to the one highlighting spillover effects from sovereign to private sector credit risk due to public debt financing of government spending. In particular, during

restrictions.

[14] Brueckner et al. (2023) use directly the residuals of a fiscal reaction function (based on data on government consumption expenditure at the NUTS-2 level) in the panel regression used to estimate fiscal multipliers which are not directly related to ERDF funding.

crisis periods, government can increase its fiscal stance to support an economy under distress using either budgetary measures (via direct spending or tax cuts) or extra-budgetary measures, via equity or state-owned bank loans injections, purchases of non-performing assets or government guarantees granted to banks and firms to address confidence concerns. Both types of fiscal policy measures rely on public debt financing (while budgetary measures have an impact on next period stock of public debt, extra-budgetary measures are contingent liabilities increasing future public debt) hence they put pressure on fiscal solvency which can spill to private sector credit risk. However, the latter channel may be more representative of national fiscal policy, where the government has the option of using deficit-financed public expenditures, which may create problems of confidence in the sustainability of the government's fiscal positions. In contrast, the financing of EU structural funds has a different mechanism, through transfers from all EU countries. Moreover, the size of the EU budget is small in relation to EU GDP and the ERDF is only a part of it. Therefore, the first channel is the one we can expect to work, thus supporting the positive effects of ERDF investments on the financial fragility of the business sector.

Thus, we construct a panel dataset of 106 Italian provinces (NUTS-3 data), over the period 2009-2018, and study the effects of the exogenous ERDF changes on the NPLs-to-output ratio, using the Local Projections approach (Jordà, 2005).

The empirical evidence we provide shows that ERDF shocks contributes to a reduction in the non-performing loans (NPLs)-to-output ratio, thereby documenting a reduction in borrowers' risk of the non-financial business sector. The most affected sectors are manufacturing and construction, while there is no impact on the NPLs of the service sector. Overall, the empirical findings show that the EU policy, funded through ERDF and mainly financing public investments, improves, as shown in the first stage of the analysis, the economic conditions of the Italian regions, with a positive effect also on the borrowers' balance sheets, thus lowering their financial fragility.

Finally, in line with Di Caro and Fratesi (2022), we account for heterogeneity. In particular, we distinguish between real (competitiveness) and financial (credit frictions) factors driving the ERDF shock impact on firms' financial fragility. For this purpose, we use Local Projections with a Smooth Transition Function to explore the presence of a non-linear relationship between the ERDF and the NPLs. We first use an index of credit supply conditions at the NUTS-1 level in Italy as transition variable in the smooth transition function. This analysis is motivated by the role played by credit supply conditions in influencing the ability of the business sector to obtain new loans. When these conditions are loosened, banks can lend more to businesses and this in turn increases the probability that some of these loans will not be repaid. Thus, the short-term counter-cyclical effects of ERDF investment on the financial fragility of the business sector, documented in the linear analysis, may be stronger in periods of credit supply easing. Indeed, we observe that the negative effects of ERDF on NPLs prevail in a regime of looser supply conditions (with a notable exception of the construction sector because the series of the index for this sector presents only one episode of credit supply easing). Second, we wonder whether ERDF investments can help less competitive Italian regions compared to more competitive ones. The degree of competitiveness is important for the business sector. Higher competitiveness indicates greater efficiency of the region's business sector and creates more opportunities for companies to grow, thus strengthening their economic conditions, with a possible lower level of financial fragility. Conversely, the business sector operating in a less competitive environment may face more difficulties. Therefore, we question whether the positive effects of the ERDF fund on the financial fragility of the business sector are greater in less competitive Italian regions. Using the Regional Competitiveness Index, and its sub-indices, as a transition variable in the smooth transition function, we find evidence that regions with a lower degree of competitiveness may benefit more from EU transfers. Interestingly, this result is most evident when looking at competitiveness in terms of innovation, which is one of the main areas of ERDF investment.

The rest of the chapter is organized as follows. Section 2.2 gives an overview of the ERDF program and reviews the empirical literature which studies the effects of the EU structural funds on the economy. Section 2.3 concerns the data used in the analysis and section 2.4 the econometric approach employed. In section 2.5 we discuss the results and finally in section 2.6 we provide some concluding remarks.

2.2 Literature review

The EU Regulation No. 1301/2013 sets out the objective of the European Regional Development Fund (ERDF). In particular, Article 5 states that it must contribute to a smart, sustainable and inclusive growth by investing in research and innovation, information and communication technologies (ICT) and small and medium-sized enterprises (SMEs). Furthermore, Article 1 states that it must contribute to the reduction of economic and structural imbalances between the regions of the EU. It is part of the European Structural and Investment Funds (ESI), whose main task is to implement EU cohesion policy. In particular, the ESI funds are intended for the implementation of EU regional policy, whose main area of intervention is economic growth, employment, sustainability, education and social equality. These relevant objectives naturally lead to conducting research to study the economic effects of EU structural funds.

Indeed, many contributions to the empirical literature on EU structural funds have focused on this aspect and studied their economic effects in terms of reducing the imbalances across EU regions and of sustaining long-run growth. A first series of works produced mixed results, in the sense that some analyses have found a positive impact on the economy (cf. Cappelen et al., 2003; Bouvet, 2005; Eggert et al., 2007), while others have found little or no impact (cf. Garcia-Milà and McGuire, 2001; de Freitag et al., 2003; Bussoletti and Esposti, 2004; Percoco, 2005; Esposti and Bussoletti, 2008; Dall'Erba and Le Gallo, 2008). These preliminary studies focus on a short time span which involves one programming period only.

More recent studies highlight the positive effects of EU structural funds on the economy. Mohl and Hagen (2009) extend the previous analyses and do not find clear results when considering the total amount of payments, but by focusing only on Objective 1 (the one that aims to help underdeveloped regions in their convergence to more developed ones), they show that EU payments contribute to regional growth. Becker et al. (2010) use a fuzzy regression discontinuity design, exploiting the eligibility rule for Objective 1 regions, which classifies the treated regions as those with a GDP per capita below 75% of the EU average. They focus on Objective 1 payments and construct a panel dataset at the NUTS-2 and NUTS-3 level, finding a positive impact on GDP growth. Pellegrini et al. (2012), instead, use a sharp regression discontinuity design, also exploiting the eligibility rule, but comparing the scenario subjected to the policy treatment with a counterfactual situation in which the policy is not implemented. Their results also show that Objective 1 payments have a positive impact on regional growth. Becker et al. (2018) extend the analysis to the new programming period and take into account the possibility of gaining or loosing Objective 1 status in a fuzzy RDD approach. Again, they find that Objective 1 regions benefit from the treatment.

Other studies, instead, concentrated on the role of EU Cohesion Policy in producing, not only long-term, but also short-term effects on the demand side, by estimating fiscal multipliers. An article by Coelho (2019) uses NUTS-2 data to estimate the EU structural fund multipliers. The author employs panel data models, and exploits the IV approach by instrumenting EU transfers using commitments as an exogenous supply of EU transfers, as these are decided long before actual spending occurs. She finds a multiplier of around 1.8 at the time of the shock, which increases to 4, three years after the shock. Espinoza and Durand (2021) consider the total ESI funds for a panel of 28 countries and estimate the effects on GDP and GVA of different sectors. They use an IV approach to identify ESI shocks by instrumenting transfers with government loans and grants. This study also shows that ESI funds have a positive impact on the economy and the results differ across sectors. An article by Canova and Pappa (2021), which our work is close to, studies the effects of the European Regional Development Fund (ERDF) and the European Social Fund (ESF) in a panel of 281 NUTS-2 regions. They estimate the multipliers region-by-region using a Bayesian instrumental variable approach. The instrument is the series of innovations in the ERDF and ESF that are not determined by macroeconomic conditions at the euro area and country level. They find that the ERDF has positive effects but is more effective in the short-term, whereas the ESF has larger effects in the medium-run.

As for Italy, a study by Aiello and Pupo (2012) estimates a growth model using panel data and finds greater positive effects of structural funds for the regions of the South than for those of the Centre-North, but this does not contribute to reducing the economic gap that persists between these two macro-areas of the country. Finally, De Stefanis et al. (2022) use a Bayesian panel VAR to estimate the European structural fund multipliers region-by-region. The shock is identified through either Cholesky factorization or GIRF. The multipliers are heterogeneous across Italian regions, higher in the South, although not statistically different from those in the Centre-North.

2.3 Data

As mentioned in section 2.1, our analysis is carried out in two stages. For this purpose, we construct two panel datasets. This section gives the details on the sources of the data and the construction of the variables employed throughout the analysis.

2.3.1 Data: first stage analysis

First, as for the analysis of the impact of the ERDF on the economy, we construct a panel dataset of 21 Italian NUTS-2 regions. The sample period is 1988-2018 due to the availability of ERDF transfers data. This is the first fund implemented among the ESI funds and therefore it has the longest time series.

It is one of the most important investment funds among the ESI and covers approximately 40% of the total budget (Canova and Pappa, 2021). Its main purpose is to strengthen the economic and structural conditions of the EU regions, especially the poorer ones, and to enable their convergence towards more developed regions. To this end, the fund invests in research and innovation, ICT, small and medium-sized enterprises (which account for a significant share of the Italian production sector) and the sustainable development of the regions. Table 2.1 presents some summary statistics on the ERDF, GDP and other indicators of the economic activity of the 21 Italian regions. The level of per capita expenditure is higher for the less developed regions of Southern Italy (e.g. Basilicata, Calabria, Campania, Molise, Apulia, Sardinia and Sicily), which have a lower level of GDP per capita and GVA per capita. This is also more evident when looking at the last column of Table 2.1, which shows the average ERDF/GDP ratio in percentage terms, where the above-mentioned regions have a higher level of this ratio than the other regions.

The data on the ERDF fund comes from the historical EU payments database by regions. The European Commission recently provided this data at the NUTS-2 level for different funds and different programming periods. Each programming period lasts 7 years but, as can be seen from the official website, there is some overlap between programming periods, because some payments planned in one programming period may be implemented in the following period. Therefore, we consider the year in which the expenditure is actually incurred, regardless of the programming period, and combine the data accordingly, to have a dataset with the total amount of expenditure during each year for each region. Moreover, since the payments are in the form of reimbursements, they are recorded after the actual expenditure has actually incurred[15], thus the EU Commission provide a modelled measure that represents the actual expenditure in each year and it is the one used in this study[16]. Hence, we use this measure as the expenditure that a region receives during the year. Since these data are in nominal terms, we use the Italian price index to obtain the measure of expenditure in real terms.

We use different proxies of the real economic activity, such as GDP, GVA and private sector GVA, all at constant prices, provided by ARDECO. We obtain a measure of private sector GVA by summing the GVA of all NACE (aggregate) sectors considered in ARDECO excluding the GVA of the non-market sector. The latter is a proxy of public sector GVA. It measures the GVA of the following subsectors: (i) public administration and defence; (ii) education; (iii) health and social work; (iv) arts, entertainment and recreation; (v) other service activities; (vi) activities of households and extra-territorial organisations and bodies. This measure is closely linked to the GVA of the general government (for a more in-depth description and a quantitative comparison between this proxy and public expenditure, see the studies by Gabriel et al., 2023 and Brueckner et al., 2023 at European level, and Cipollini and Frangiamore, 2023, along with the first chapter of this thesis, for Italy).

[15]The way data is recorded makes ERDF spending more sensitive to the business cycle conditions of each region and thus more endogenous, further invalidating the use of Cholesky identification.

[16]For a detailed explanation of the model based public spending financed through EU funds, see the official website.

Table 2.1: ERDF and real economic activity data, average by NUTS-2 regions 1988-2018 (percapita).

NUTS-2 Regions	ERDF	GDP	GVA	GVA private sector	ERDF % of GDP
Abruzzo	43.47	24438.62	21968.62	17047.58	0.18
Basilicata	139.82	19648.05	17662.23	13442.69	0.70
Calabria	120.77	16401.41	14743.73	10396.29	0.73
Campania	92.50	18539.42	16665.65	12356.53	0.49
Emilia-Romagna	4.44	32968.31	29636.22	25197.09	0.01
Friuli-Venezia Giulia	10.05	29151.07	26204.78	21064.35	0.03
Lazio	7.78	33989.05	30553.79	23727.81	0.02
Liguria	17.55	30310.40	27246.94	22037.74	0.06
Lombardia	2.64	35913.99	32284.18	28301.00	0.01
Marche	14.06	26546.01	23863.02	19736.68	0.05
Molise	98.60	21054.69	18926.72	13681.59	0.46
Piemonte	14.21	29084.97	26145.37	22108.61	0.05
Provincia Autonoma di Bolzano/Bozen	9.70	39329.99	35354.93	28550.15	0.02
Provincia Autonoma di Trento	7.14	36055.16	32411.09	25930.15	0.02
Puglia	88.21	18270.44	16423.85	12407.96	0.48
Sardegna	92.28	19785.14	17785.47	12609.94	0.45
Sicilia	95.28	17762.80	15967.52	11235.41	0.53
Toscana	13.53	29151.01	26204.73	21664.58	0.05
Umbria	29.35	26389.05	23721.92	18825.89	0.11
Valle d'Aosta/Vallée d'Aoste	29.89	37259.16	33493.40	24945.46	0.08
Veneto	7.69	30610.25	27516.48	23814.99	0.02

2.3.2 Data: second stage analysis

We construct a panel dataset of 106 Italian provinces over the period 2009-2018, collecting data on the stock of NPLs from the Bank of Italy[17]. The data are recorded quarterly, thus to match the annual frequency of the ERDF data we take the average of the quarters of each year to obtain annual observations on non-performing loans. We consider NPLs of non-financial corporations (including the disaggregated data for the manufacturing, construction and services sector; see Table 2.2 for summary statistics)[18].

Table 2.2: Percentage share of real non-performing loans in real GDP by NUTS-1 area (averages 2009-2018).

NUTS-1	NFC	NFC (manufacturing)	NFC (construction)	NFC (services)
Northwest	4.73	1.34	1.24	2.10
Northeast	5.66	1.66	1.69	2.23
Centre	6.81	1.65	1.85	3.16
South	4.99	1.53	1.13	2.16
Islands	4.48	0.97	1.06	2.17

As for the analysis of the non-linear effects of ERDF shocks on NPLs, we employ two indicators as switching variables. We use a diffusion index, provided by the Bank of Italy at macro-regional level, to proxy credit supply conditions (taking positive values in periods of loan supply tightening and negative ones in periods of loan supply easing) and we use it as a transition variable to capture non-linear effects of

[17]We use the table TRI30211 of the BDS database (Base Dati Statistica).

[18]We do not transform this variable into real terms because the econometric approach we use allows us to control for inflation by including time fixed effects (see section 2.3). However, we repeat the analysis by transforming non-performing loans into real terms, using the Italian price index, only for the total NFCs, and the results do not change (results available upon request).

ERDF shocks in the panel Local Projections with a Smooth Transition Function. This index is constructed on the results of the Regional Bank Lending Survey, and it is available at semi-annual frequency from 2009 onwards at the macro-area level (Northwest, Northeast, Centre, South and Islands)[19]. Finally, for the analysis about the non-linearity driven by the degree of competitiveness, we use the Regional Competitiveness Index, provided by the EU Commission at the NUTS-2 level. The index is constructed using variables and indicators representing 11 pillars, grouped into the following sub-indexes: (i) Basic sub-index includes Institutions, Macroeconomic Stability, Infrastructures, Health and Quality of Primary and Secondary Education; (ii) Efficiency sub-index includes Higher Education, Training and Lifelong Learning, Labour Market Efficiency, and Market Size; (iii) Innovation sub-index includes Technological Readiness, Business Sophistication, and Innovation (Dijkstra et al., 2011). The index is updated every three-year on the basis of the past performance across different pillars. Therefore, we use the 2013, 2016 and 2019 data for the transition variable observed over the time span 2009-2011, 2012-2015, and 2016-2018, respectively. We do this because the raw data used to construct the index in each round are relative to these respective periods, in the sense that, for example, for the construction of the index in the 2019 round, they used variables related to the period 2016-2018. Moreover, the competitiveness of a region is a structural factor that can hardly change in the short-run[20].

2.4 Econometric strategy

As we have already mentioned, our strategy is based on two stages. This section discusses the econometric strategy employed in each of the two. In the first one, we identify the ERDF shocks and estimate the multipliers for some real economic activity variables. In the second one, we analyze the impact of the ERDF shock on the credit market of the Italian provinces during the period 2009-2018, mostly characterized by financial and sovereign debt crises.

2.4.1 Empirical analysis: identification of ERDF shocks

In the first step, we use a SVAR approach to study the effects of the ERDF fund on the economy and to compute the multipliers. The VAR model was introduced by Sims (1980), and allows to treat all the variables as endogenous.

The data are, first, transformed following the suggestions of Ramey and Zubairy (2018), hence they are scaled by potential output, using the Hodrick-Prescott filter (setting the smoothing parameter equal to 100 for annual data) fitted to GDP (or total and private GVA, when the focus is on the multipliers of this proxy of real economic activity)[21]. Then, we fit to the transformed data the following reduced-form panel VAR(1)[22]:

$$Y_{i,t} = \alpha_i + \gamma_t + AY_{i,t-1} + u_{i,t} \tag{2.1}$$

[19] It corresponds to the NUTS-1 classification.

[20] In an exercise, we also use a dummy variable in the Local Projections, taking 1 if the region is recorded to have a value of the Regional Competitiveness Index below the Italian median. The results do not change.

[21] Ramey and Zubairy (2018) argue that the alternative method, based on fitting the VAR model data to log-levels, and then convert the elasticities into multipliers using a scaling factor calculated as the ratio of the sample mean of GDP to government expenditure, implies very sensitive estimates of the multipliers to the sample period considered.

[22] We follow the Stock and Watson criterion to set the number of lags. We include one lag because our dataset is at annual frequency.

where $Y_{i,t} = [ERDF_{i,t}, GDP_{i,t}]$, is a vector of endogenous variables containing ERDF and GDP, both in real terms, in region i at time t, divided by the GDP trend. For robustness, we also consider as a proxy of output, the GVA and the GVA of the private sector scaled by the corresponding trend. $Y_{i,t-1}$ is the vector containing the lagged endogenous variables, A is a 2×2 matrix of coefficients of the lagged endogenous variables, and α_i and γ_t are, respectively, regional and time fixed effects. The vector of the two reduced-form innovations is $u_{i,t}$ which relates to the structural shocks according to the so-called B-form representation (Lütkepohl, 2005), as in the following equation:

$$u_{i,t} = B\epsilon_{i,t} \tag{2.2}$$

where $\epsilon_{i,t}$ is the vector of structural shocks, and the matrix B contains the structural coefficients, measuring the impact of the structural shocks on the variables of the system. The structural form is therefore as follows:

$$Y_{i,t} = \alpha_i + \gamma_t + AY_{i,t-1} + B\epsilon_{i,t} \tag{2.3}$$

As outlined in the main introduction of the thesis, we estimate a pooled panel VAR with regional and time fixed effects following the literature that focuses on the estimation of the "geographic cross-sectional fiscal spending multiplier" (see Nakamura and Steinsson, 2014 and Chodorow-Reich, 2019). According to Nakamura and Steinsson (2014), this approach, together with the use of time fixed effects, allows controlling for common shocks, such as monetary policy, which is a major confounding factor in this type of analysis. This is an important advantage when estimating the multiplier in this context. Moreover, the panel time series dimension may not be long enough to apply the mean group estimator of Pesaran and Smith (1995)[23]. We also recognise that there are important issues related to spillover effects between regions. This type of analysis deserves special attention especially when the sample size T, regarding the time series dimension of the panel, is short. Canova and Pappa (2007) argue that the shortness of the dataset may prevent the use of richer models to study the transmission of shocks between units. This essentially motivates the choice of constraining the parameters in equation (2.3) to be homogeneous across regions.

Identification. We acknowledge the dependence of fiscal spending financed by ERDF on regional economic conditions and, to retrieve the exogenous component of public spending, we address the issue of identification of fiscal shocks by using a panel proxy-SVAR. The identification scheme applied to structural VAR was first introduced by Stock and Watson (2012) and Mertens and Ravn (2013), and extended by Gertler and Karadi (2015). To our knowledge, our study is the first to extend the methodology (in particular the Gertler and Karadi, 2015 approach) to a panel framework. More specifically, we denote with $Y^p_{i,t}$ the policy indicator observed in region i at time t, which in our case is the ERDF, with an associated structural shock $\epsilon^p_{i,t}$. The structural form we wish to estimate is as follows:

[23] However, we are aware of the availability of approaches that allow some form of partial pooling. This is an area that we are still exploring and we leave this for future research.

$$Y_{i,t} = \alpha_i + \gamma_t + AY_{i,t-1} + B_1\epsilon^p_{i,t} \tag{2.4}$$

where B_1 is the first column of the B matrix, containing the impact response of the endogenous variables to an ERDF shock. We need an instrument $Z_{i,t}$ for the policy shock to identify the first column of matrix B. The instrument should satisfy two conditions, i.e., it must be correlated with the policy shock $\epsilon^p_{i,t}$, and orthogonal to the non-policy shock $\epsilon^q_{i,t}$, that is:

$$E[Z_{i,t}\epsilon^{p'}_{i,t}] = \phi \tag{2.5}$$

$$E[Z_{i,t}\epsilon^{q'}_{i,t}] = 0 \tag{2.6}$$

To obtain such an instrument for the ERDF shock, we need a measure that is not affected by the state of the regional economies. We follow Canova and Pappa (2021) and Brueckner et al. (2023) in doing this[24]. Canova and Pappa (2021) estimate a fiscal reaction function using Euro Area variables and country-level data, and instrument EU regional transfers through its innovations to Euro Area and country-level macroeconomic variables, obtained from the residual of the above mentioned fiscal reaction function. They use this instrument in a dynamic single panel equation. Instead, we estimate the fiscal reaction function at the NUTS-2 level, to obtain a series of ERDF expenditure orthogonal to current and past regional economic conditions. Moreover, Brueckner et al. (2023) estimate a similar fiscal reaction function for nationally-financed government consumption expenditure, but they use the residuals of this model directly as regressor in a dynamic single panel equation of output on exogenous government spending, represented by such residuals. Although the strategy we adopted is broadly similar to the one used by Brueckner et al. (2023) and Canova and Pappa (2021), we followed the IV local projection approach of Canova and Pappa (2021), which is equivalent to a proxy SVAR context based on the direct approach, hence by directly instrumenting the target shock with the residuals of the fiscal reaction function (see below for the details about the estimation procedure).

More specifically, we generate a series of ERDF changes that are orthogonal to GDP by taking the residuals estimated from the following panel equation:

$$ERDF_{i,t} = \alpha_i + \gamma_t + \beta_1 GDP_{i,t} + \beta_2 ERDF_{i,t-1} + \beta_3 GDP_{i,t-1} + r_{i,t} \tag{2.7}$$

where α_i and γ_t are regional and time fixed effects. We account for reverse causality of GDP by using an instrumental variable approach to estimate equation (2.7) (Fatás and Mihov, 2013; Alesina et al., 2008). Following Alesina et al. (2008), we instrument the GDP of region i using the GDP of the other regions belonging to the same NUTS-1 area as region i, excluding the GDP of region i. This ensures that the GDP of the other regions in the same NUTS-1 area is somehow related to the GDP of region i, satisfying the relevance condition. Furthermore, the exogeneity assumption is based on the fact that using the GDP of the NUTS-1 area to which region i belongs, minus the GDP of region i, ensures that this IV is such

[24]Fatás and Mihov (2013) and Alesina et al. (2008) estimate a similar model for government spending in the context of fiscal policy pro-ciclicality and policy volatility.

that it is not affected by variations in ERDF transfers to region i that affect the GDP of region i, i.e., the IV is not contaminated by the region-specific GDP that is the source of endogeneity[25].

We test the relevance of the IV for GDP by computing the F-test of the first stage regression. We use robust estimators, especially the Driscoll and Kraay (1998) estimator which takes into account the heteroskedasticity, serial correlation and correlation across units, and the Kleibergen-Paap rk Wald F-statistic, which, as showed by Andrews et al. (2019), is equivalent to a non-homoskedasticity-robust-F-statistic to test the relevance of the IV, in the case of one endogenous variable and one instrument, and this statistic must be compared with the critical values of Stock and Yogo (2005)[26]. The results in Table 2.3 show that the F-statistic is above the Stock et al. (2002)' threshold of 10 and the K-P statistic is greater than the critical value for the 15% maximal IV size. This means that our instrument, constructed as described above, satisfies the relevance condition for regional GDP[27].

Table 2.3: First stage results for equation (2.7).

	coeff.	std. errors
$IV_{i,t}$	0.50***	(0.14)
$GDP_{i,t-1}$	0.35***	(0.13)
$ERDF_{i,t-1}$	0.61#	(0.36)
F-statistic IV		12.25
Kleibergen-Paap stat.		10.75

Note: Driscoll and Kraay (1998) robust std. errors and F-stat. for IV. *** and # indicate statistical significance at 0.1 and 10% levels.

Table 2.4 shows the results of the IV estimation of equation (2.7). We find that EU transfers were counter-cyclical in the period of our study, the sign of the output coefficient being negative, although it is not statistically significant[28]. The table also shows that past ERDF spending has a significant effect on future spending, while future spending is not affected by the past value of output.

Table 2.4: Results of IV estimation of equation (2.7).

	coeff.	std. errors	t value	p-value
$GDP_{i,t}$	-0.0158025	0.0125831	-1.2558	0.2097
$GDP_{i,t-1}$	0.0044395	0.0051334	0.8648	0.3875
$ERDF_{i,t-1}$	0.6961285	0.1041421	6.6844	5.478e-11 ***

Note: Driscoll and Kraay (1998) robust std. errors. *** indicates statistical significance at 0.1 level.

The panel structure of our dataset, using regional and time fixed effects, allows to control for confounding factors. More specifically, the inclusion of the regional fixed effects is crucial because this allows us to control for the different structural conditions[29] among the regions, that characterize the North-South

[25] Alesina et al. (2008) constructed the IV for the output gap in a panel of countries by aggregating the output gap of the area in which the country is located, excluding the output gap of this country. We follow this idea to construct the IV for Italian regional outputs. However, we must recognise that the exogeneity hypothesis may become weaker in the presence of spillover effects produced by ERDF spending, dynamic GDP interrelationships and in general cross-border economic interactions between Italian regions.

[26] 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; (iv) for 25% maximal IV size is 5.53.

[27] When estimating the effects on total GVA and private sector GVA, we replace GDP with each of them at a time in equation (2.7) and construct the IV in the same way, but using total GVA and private sector GVA. The robust F-statistic of Driscoll and Kraay (1998) is 12.15 for the IV of total GVA and 16 for the IV of private sector GVA, while the K-P statistic is 10.75 for the IV of total GVA and 8.57 for the IV of private sector GVA.

[28] This, of course, can not be considered as a test of the exogeneity of the ERDF expenditure to output within a year, to justify the use of Cholesky restrictions.

[29] This is because of the fact that structural conditions may be quite stable over time and in the peculiar case of Italy, the North-South divide is a centuries-old issue and still persists today.

divide of Italy. Since the ERDF aims at reducing the imbalances among the regions, if we do not include regional fixed effects, the correlation between ERDF transfers and region-specific characteristics may bias the results. Moreover, the inclusion of time fixed effects allows us to control for aggregate common shocks, such as the ECB monetary policy (see Gabriel et al. 2023). We argue that the use of time fixed effects can also control for national tax policy, given that, mainly, tax policy in Italy is decided by the central government, implying that public spending at the local level is not directly related to revenues, due to central government transfers acting as the main source of finance. Moreover, in the extent to which tax policy is not common across regions, as argued by Coelho (2019), it may not represent a confounding factor in the context of the EU structural funds for two order of reasons: (i) public revenues, that support the EU spending, are decided during supranational negotiations, much before the actual spending exert its effect on the regional economies; (ii) the EU structural spending is financed using contributions from the member states to the overall EU budget. This budget is dedicated to financing programs other than those financed by the ERDF, thus these contributions are not directly related to ERDF expenditure.

Estimation. First, we estimate equation (2.7), and take the residual as a series of changes in the ERDF that are orthogonal to the non-policy shock, which means $Z_{i,t} = \hat{r}_{i,t}$. Then, we estimate the model in equation (2.4), following the steps below:

- Estimate the reduced-form VAR from equation (2.1) and take the estimated innovations $u_{i,t}$[30];
- Let $u^p_{i,t}$ be the reduced-form residual of the equation for the policy indicator. Regress $u^p_{i,t}$ on the instrument $Z_{i,t}$ (first stage) and get the fitted values $\hat{u}^p_{i,t}$ from this first stage regression, that will represent the ERDF expenditure not driven by regional output;
- Let $u^q_{i,t}$ be the reduced-form residual of the equation for the GDP. To get an estimate of the response of the GDP to an ERDF shock, regress $u^q_{i,t}$ on $\hat{u}^p_{i,t}$ and normalize the on impact response of the ERDF to an its own shock to one. In this way, up to this normalization, which means $B_{1,1} = 1$, we identify the first column of the matrix B, where $B_{2,1}$ is the coefficient of the regression of $u^q_{i,t}$ on $\hat{u}^p_{i,t}$:
$$u^q_{i,t} = \frac{B_{2,1}}{B_{1,1}}\hat{u}^p_{i,t} + \psi_{i,t} \tag{2.8}$$
- Collect the on impact estimated responses in $B_1 = [1, B_{2,1}]^{'}$.

In equation (2.8), $\hat{u}^p_{i,t}$ is orthogonal to $\psi_{i,t}$ given the assumption (2.6). Then, we can compute the Impulse Response Functions to an ERDF shock occurring at time t as follows:

$$IRF(h) = A^h B_1 \tag{2.9}$$

The first and second elements of the vector $IRF(h)$ measure, respectively, the response of the ERDF and GDP in $t+h$ to an ERDF shock in t[31].

[30] We estimate the reduced-form VAR with panel data, applying the OLS estimator on within transformed data, to control for regional and time fixed effects. Even though each equation in the VAR is a dynamic panel data model, the time series dimension of our panel dataset, equal to 31, is high enough to avoid the Nickell (1981) bias, as suggested by Monte-Carlo evidence in Judson and Owen (1999).

[31] The procedure described in the three steps leading to eq (2.8) turns out to give the same results if I had used an indirect

To compute the multipliers, we follow the suggestion of Ramey and Zubairy (2018), and calculate the multiplier at horizon h as the cumulated response of GDP divided by the cumulated response of the ERDF:

$$m(h) = \frac{\sum_{j=0}^{h} IRF_j^{GDP}}{\sum_{j=0}^{h} IRF_j^{ERDF}} \tag{2.10}$$

With regard to the inference, we construct bootstrap distributions for the IRFs in order to obtain confidence interval bands. We use the pairs bootstrap for panel data, as in Kapetanios (2008), resampling the observations in the time dimension for each region, which means creating a bootstrap dataset by randomly extracting with replacement one year for each region and forming the pairs $[Y_t^*, Y_{t-1}^*]$, where $Y_t^* = (y_{1,t}^*, y_{2,t}^*, ..., y_{N,t}^*)$ and $Y_{t-1}^* = (y_{1,t-1}^*, y_{2,t-1}^*, ..., y_{N,t-1}^*)$. Therefore, for each repetition we resample the dataset and estimate everything, including equation (2.7). We perform 1000 repetitions and calculate the 16th and 84th percentiles of the bootstrapped IRF distributions to get the 68% confidence interval bands.

2.5 Empirical analysis: ERDF shocks and financial fragility of the private sector

In the second stage of the analysis, we investigate the effects of the shock to ERDF funds on the financial fragility of the private sector, using provincial data for non-performing loans available since 2009. We transform the volume of non-performing loans as a share of potential GDP in order to use the same transformation of the variables employed in the VAR, and we use the Local Projections approach, developed by Jordà (2005), to estimate the effects of the ERDF shock (retrieved from the first stage of the analysis) on the ratio of non-performing loans to potential GDP. In particular, for each horizon h, we estimate the following panel regression:

$$y_{j,t+h} = \alpha_{h,j} + \gamma_{h,t} + \beta_h \hat{u}_{i,t}^p + \delta_h X_{i,j,t-1} + \nu_{i,t} \tag{2.11}$$

where $\alpha_{h,j}$ and $\gamma_{h,t}$ are provincial and time fixed effects[32], respectively, $X_{i,j,t-1}$ is a vector of control variables at the NUTS-2 and NUTS-3 level, in particular we include a lag of $\hat{u}_{i,t}^p$, a lag of $y_{j,t}$ and a lag of real GDP divided by potential GDP at the provincial level[33]. Given the panel structure used in this analysis, which has $N = 106$ and $T = 10$, the fixed effects estimator is biased, due to the presence of the lag of the dependent variable among the regressors (Nickell, 1981). Therefore, we address this issue by estimating the parameters of equation (2.11) using the GMM-sys approach, developed by Blundell and Bond (1998), which consists of extending the GMM-diff of Arellano and Bond (1991), adding the equations in levels, and instrumenting the lags of the endogenous variables in levels using appropriate

approach within the proxy SVAR framework (Caldara and Kamps 2017 and Angelini et al. 2024), identifying the ERDF shock by instrumenting the output shock, instead of estimating the reaction function in equation (2.7). I modified the code, provided in the link attached at the end of the thesis, according to the suggestions provided by the referee, Prof. Luca Fanelli, and I obtained the same results. I am grateful to him for this useful clarification. This also holds for Chapter 3, where I used the same methodology.

[32] Again, similar to the case of the model in section 2.4.1, the inclusion of provincial and time fixed effects allows us to control for time-invariant provincial characteristics and aggregate common shocks.

[33] In this model, j indexes the province and i the region.

lags of the first differences of these endogenous variables. Since the time series is short, we make use of all available instruments.

The Local Projections approach allows to estimate the impulse response function at horizon h by taking the estimated coefficients $\hat{\beta}_h$ from equation (2.11). Therefore, the series of $\hat{\beta}_h$'s will deliver an estimate of the dynamic response of the NPLs-to-output ratio to a regional shock hitting the ERDF expenditure.

We, further, extend the model given by (2.11) to account for non-linearity using the following panel Smooth Transition Regression (see Auerbach and Gorodnichenko, 2013, among the others):

$$\begin{aligned} y_{j,t+h} = & (1 - F(z_{k,t})) \left[\alpha_{h,j}^{(1)} + \gamma_{h,t}^{(1)} + \beta_h^{(1)} \hat{u}_{i,t}^p + \delta_h^{(1)} X_{i,j,t-1} \right] \\ & + F(z_{k,t}) \left[\alpha_{h,j}^{(2)} + \gamma_{h,t}^{(2)} + \beta_h^{(2)} \hat{u}_{i,t}^p + \delta_h^{(2)} X_{i,j,t-1} \right] + \nu_{i,t} \end{aligned} \tag{2.12}$$

where the superscripts (1) and (2) indicate the regimes. Therefore, $\beta_h^{(1)}$ gives the point estimate of the IRF at horizon h in the first regime, whereas $\beta_h^{(2)}$ in the second one. The specification of the smooth transition function is as follows:

$$F(z_{k,t}) = \frac{e^{-\gamma z_{k,t}}}{1 + e^{-\gamma z_{k,t}}} \tag{2.13}$$

where γ is the smoothing parameter that governs the shape of the function (the higher it is, the more abrupt is the change in the regime) and $z_{k,t}$ is the transition variable. As mentioned in section 2.3.2, we use two different transition variables. To study the effects of ERDF shocks on the financial fragility of the business sector, conditional on credit supply conditions, we use the credit supply diffusion index. This index is constructed by the Bank of Italy based on the results of the Bank Lending Survey. The index takes positive values in periods of credit supply tightening, while it is negative in periods of credit supply easing; therefore, the higher the index, the tighter the credit supply conditions. Since the smooth transition function, $F(z_{k,t})$, is a decreasing function of the transition variable, this means that the tighter the credit supply conditions, the lower the value of this function, which in turn means that the first regime in equation (2.12) is that of tight supply conditions, while the second is that of looser conditions. The second transition variable is the Regional Competitiveness Index, provided by the European Commission. The higher the index, the more competitive the region. Therefore, since $F(z_{k,t})$ is a decreasing function of this index, as soon as competitiveness increases, the smooth transition function tends to zero, so the first regime is the one characterised by a high degree of competitiveness while the second by a low degree of competitiveness. These indicators are standardized before entering the smooth transition function[34].

2.6 Empirical evidence

The results are discussed in the next sub-sections. First, we show the results of the first stage, obtained by estimating the SVAR model, and then we present the results obtained in the second stage about the

[34] We calibrate the γ parameter at the value of 5 to allow for an intermediate degree of intensity in the regime switching (Colombo et al., 2022), but the results remain roughly the same qualitatively for other values of this parameter (results available upon request).

credit market.

2.6.1 Empirical evidence: first stage analysis

We, first, test the relevance condition, given by eq. (2.5), as in Gertler and Karadi (2015) and Caldara and Kamps (2017), by regressing the policy reduced-form residual $u^{p}_{i,t}$ (the residual estimated in the first VAR equation, which is the one for the ERDF equation) on the proxy $Z_{i,t}$, and take the robust F-test, computed using the HAC Newey-West estimator. We find an F-test of 3632.32 in the model for GDP, 3612.26 in the model for GVA and 12309.33 in the model for private sector GVA. These values are all above the threshold of 10 suggested by Stock et al. (2002), thus showing that the proxies satisfy the condition (2.5).

The results for the elasticity of the GDP, GVA and private sector GVA to ERDF shocks are given by the Impulse Response Functions shown in Figures 2.1, 2.2 and 2.3, together with the 68% confidence bands (dashed lines).

Figure 2.1: Impulse Response Functions and 68% confidence bands (GDP).

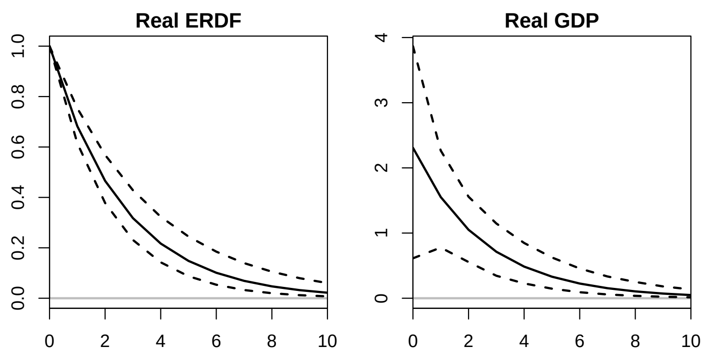
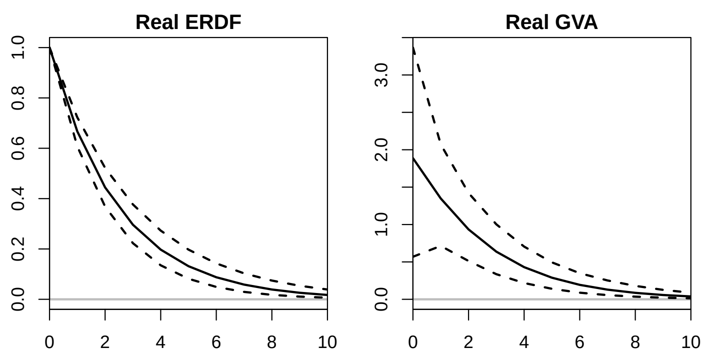
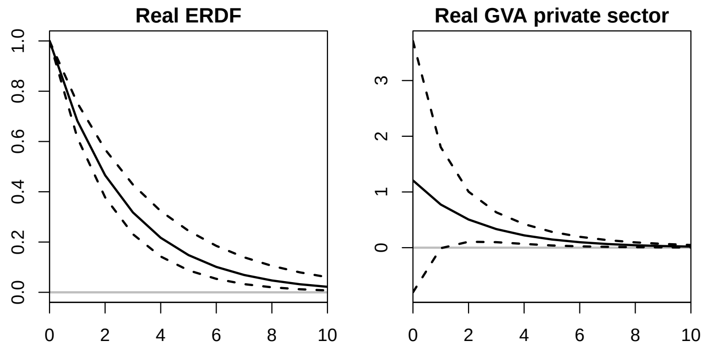


Figure 2.2: Impulse Response Functions and 68% confidence bands (GVA).



From Figure 2.1 we can observe that both ERDF and GDP increase significantly after a positive ERDF shock and that the response decreases over time and vanishes about seven years after the shock. The response of total GVA is slightly lower than that of GDP, but it follows a similar path (see Figure 2.2). In contrast, the response of private sector GVA is significant from 2 years after the shock and is lower than that of GDP and total GVA (see Figure 2.3). The estimated multipliers are shown in Table 2.5. We find multipliers for GDP ranging between 2.31 and 2.28, respectively on impact and 3 years

Figure 2.3: Impulse Response Functions and 68% confidence bands (private sector GVA).



after the shock, for total GVA between 1.89 and 2, respectively on impact and 3 years after the shock, and, for private sector GVA, 1.18 and 1.17, 2 and 3 years after the shock, respectively. Notably, the latter two results are close to the ERDF multipliers found by Canova and Pappa (2021) for private sector GVA. Therefore we find positive short-term effects of the ERDF on the economies of the Italian regions. Furthermore, the empirical evidence shows multipliers for the ERDF above one. This is consistent with the literature finding that government investment multipliers are above the unity (see for instance Auerbach and Gorodnichenko, 2012 for the U.S. and Deleidi et al., 2021 for Italy).

Table 2.5: Multipliers.

years	0	1	2	3
GDP	**2.31**	**2.29**	**2.29**	**2.28**
GVA	**1.89**	**1.94**	**1.98**	**2.00**
private GVA	1.21	1.19	**1.18**	**1.17**

Note: in bold are the statistical significant multipliers, based on the significance of IRFs.

2.6.2 Empirical evidence: second stage analysis

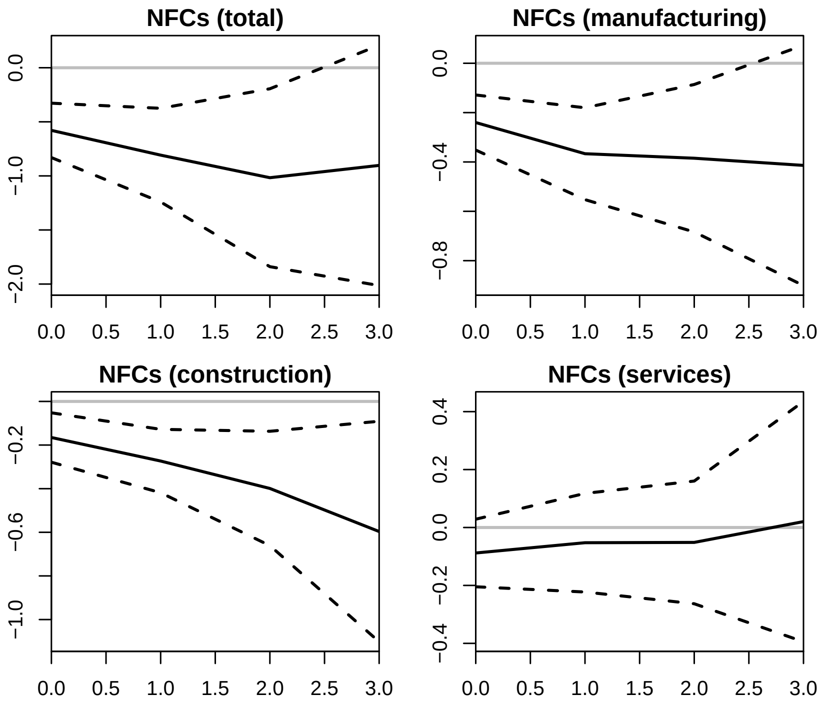
Linear response of NPLs. Figure 2.4 shows the linear impulse response plot (with the 95% confidence bands represented by dashed lines) of the non-performing loans, NPL (e.g. the ratio to potential GDP), used as a proxy for the financial fragility of the private sector, through Local Projections (see eq. 2.11).

As already mentioned, we study the short-to-medium-term impact of the ERDF shock on the NPLs-to-output ratio of the non-financial corporate sector as a whole and in its decomposition into the manufacturing, construction and service sectors.

First, we can observe that, overall, an increase in the ERDF significantly reduces the NPLs-to-output ratio. Overall, this improvement of NPLs for non-financial corporations can be explained by acknowledging that ERDF finance investments in projects that can benefit businesses (in particular, ERDF finance investments in Research and Development, ICT and SMEs). Moreover, if we look at the breakdown of NFCs into the different sectors, we find that those that decrease the most are the NPLs of the manufacturing and construction sectors, while those of the service sector do not react. Indeed, the ERDF should be considered as investment in the manufacturing and research, because of the aim of the fund and the area of intervention (Canova and Pappa, 2021).

We argue that our empirical evidence is consistent with the credit channel discussed in Auerbach et al.

Figure 2.4: Non-performing loans Impulse Response Functions by borrowers and sectors, and 95% confidence bands.



(2020) and it shows that public investment spending, channeled through the ERDF fund, by improving the economic conditions of the regions, contributes to reducing the risk of borrowers and, consequently, to lowering non-performing loans. Overall, the empirical evidence shows that the positive short-term effects of the ERDF on the economy plays a role in dampening the consequences of the financial crises, occurred over the period 2009-2018, on the financial fragility of the non-financial corporations.

Non-linear response of NPLs. We now investigate whether there is evidence of a heterogeneous response of NPLs to an ERDF shock. In line with Di Caro and Fratesi (2022), we acknowledge that economic and structural differences across regions can shape the way in which regions react to shocks. Therefore, we study whether the benefits for the firms' financial fragility coming from the ERDF investment, during the financial crises, are different by conditioning on financial (credit frictions) and real (competitiveness) factors, characterizing the Italian regions.

First, the estimation of the Smooth Transition Local Projections (see equation 2.12), using the credit supply diffusion index as transition variable, shows (see Figure 2.5), for the credit supply easing regime, a decrease of the NPLs-to-output ratio in response to an increase in ERDF spending, especially for the manufacturing sector[35]. Consequently, fiscal spending through ERDF does not reduce the borrowers' credit risk during periods associated with binding credit constraints, but it does in period of credit supply easing, when it is more likely that banks provide loans, and in particular, the credit obtained by riskier borrowers may increase.

Finally, we examine whether the effects of the ERDF on the NPLs can be different across regions

[35] It is important to note that there is only one episode of credit supply easing for the construction sector, implying almost no distinction between the regimes.

characterized by different degree of competitiveness. The empirical evidence in Figure 2.6 shows that fiscal spending funded by ERDF significantly lowers the NPLs of the business sector (mainly the manufacturing and construction sectors) in regions with a low degree of competitiveness (measured by the Regional Competitiveness Index). When we turn our focus on the three main different pillar of the Regional Competitiveness Index (see Figures 2.7, 2.8 and 2.9) we can observe that the financial fragility improvement (due to an increase in ERDF) in the manufacturing sector is associated to the basic and innovation sub-indices, while the positive effect for construction can be related only to the innovation sub-index. The positive role played by the innovation pillar of the Regional Competitiveness Index can be explained by the main features of ERDF funds tailored to fund projects with high innovative content (in the area of research and development and those related to information and communication technologies). Consequently, the empirical evidence suggests that fiscal spending through ERDF reduces territorial disparities related to borrowers credit risk driven by a low degree of regional competitiveness.

Figure 2.5: Non-performing loans Impulse Response Functions (business sector) under tight (dark gray) and slack (light gray) credit supply conditions regimes, with 95% confidence bands.

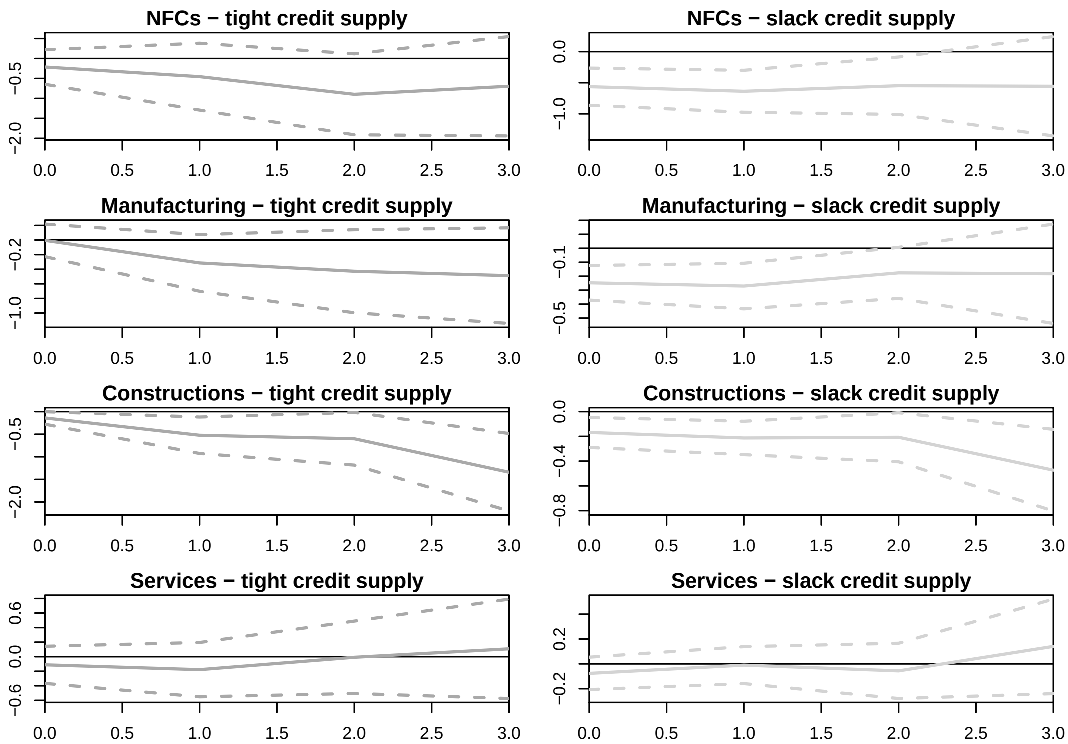


Figure 2.6: Non-performing loans Impulse Response Functions (business sector) under high (dark gray) and low (light gray) general competitiveness regimes, with 95% confidence bands.

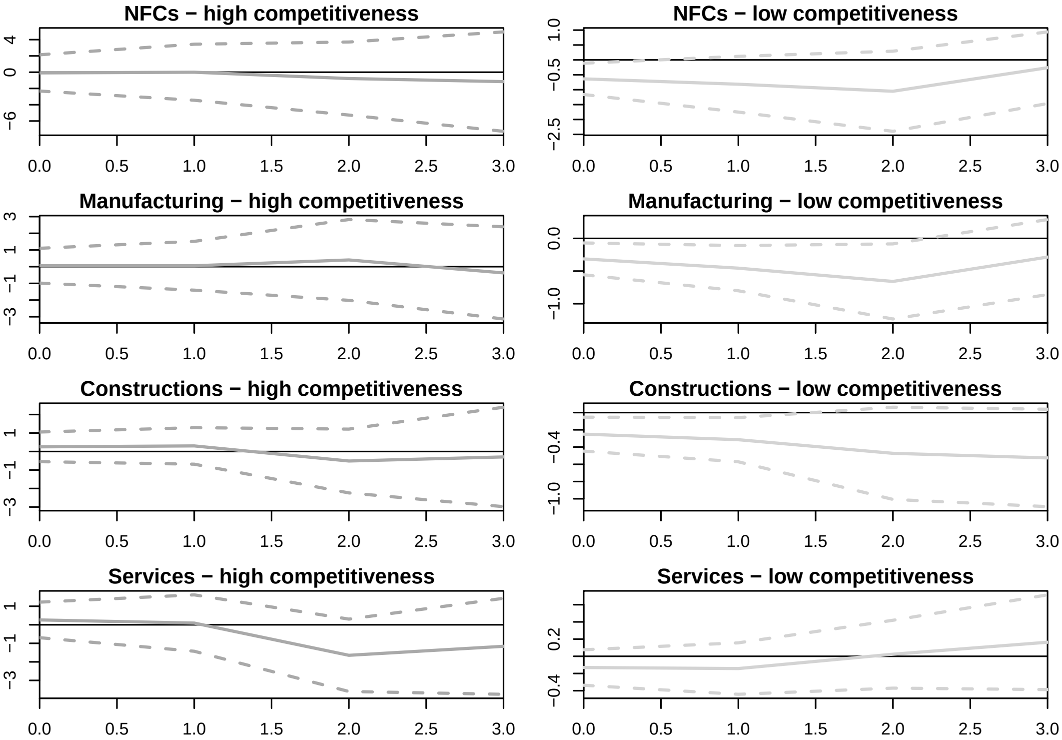


Figure 2.7: Non-performing loans Impulse Response Functions (business sector) under high (dark gray) and low (light gray) innovation competitiveness regimes, with 95% confidence bands.

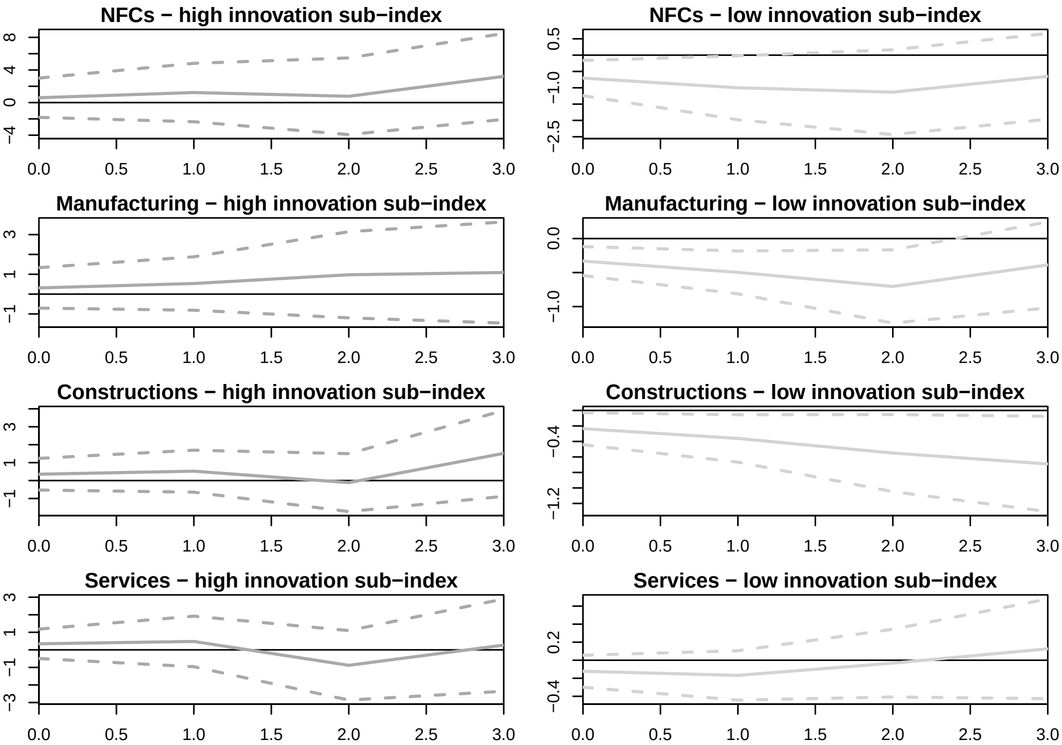


Figure 2.8: Non-performing loans Impulse Response Functions (business sector) under high (dark gray) and low (light gray) basic competitiveness regimes, with 95% confidence bands.

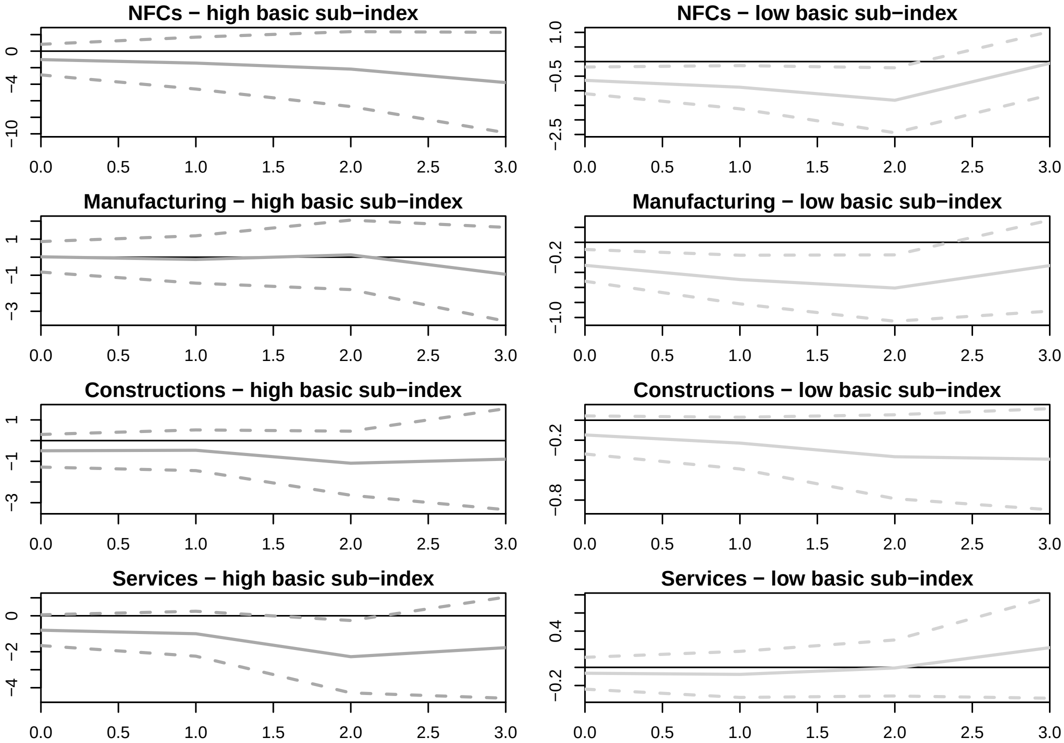
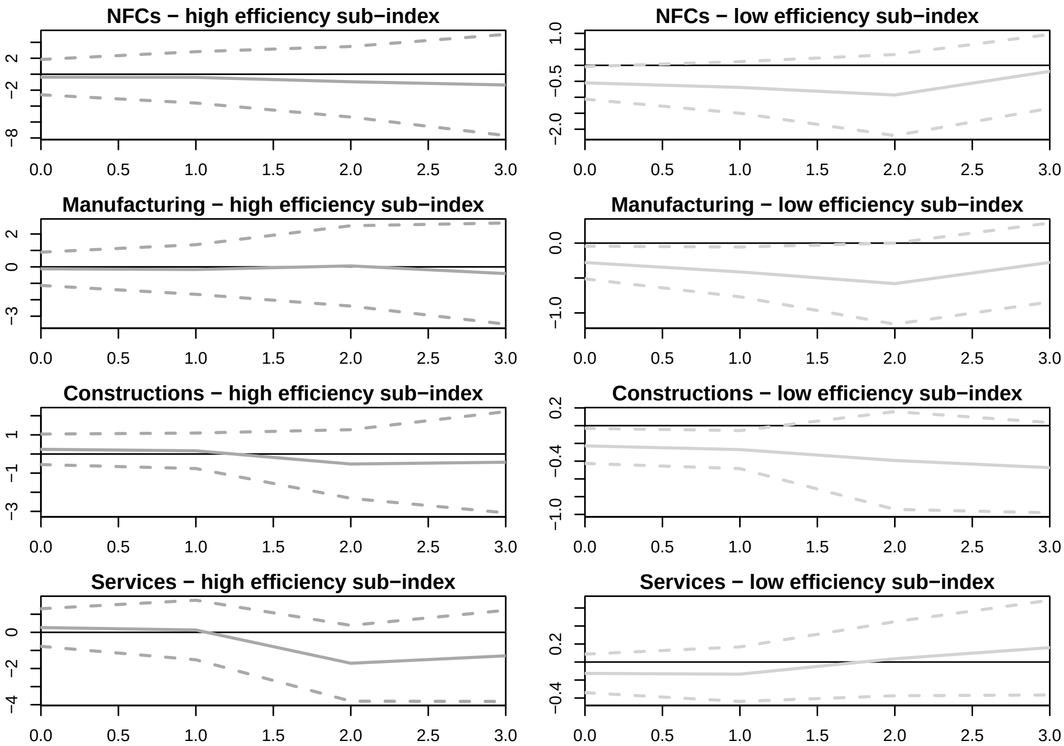


Figure 2.9: Non-performing loans Impulse Response Functions (business sector) under high (dark gray) and low (light gray) efficiency competitiveness regimes, with 95% confidence bands.



2.7 Conclusions

In this study we examine whether public investments financed through the European Regional Development Fund, contributed to reduce the financial fragility of firms in Italy, during the 2009-2018 sample period, mostly characterized by episodes of real and financial turmoil. In particular, our focus is on the exogenous component of ERDF expenditure and this requires to resort to the identification of shocks to ERDF.

Therefore, we, first, study the impact on the economy, exploiting a panel of 21 Italian NUTS-2 regions over the period 1988-2018. We use a SVAR approach and identify ERDF shocks by implementing the proxy-SVAR, where the instrument is an estimated series of ERDF changes orthogonal to the economic activity of the regions. Furthermore, we analyse the effects of ERDF shocks on the credit market, in particular on the non-performing loans-to-potential output ratio, to investigate whether the positive impact of the ERDF spending on the economy, contributes to reducing the probability of a local recession and, in turn, to lowering the risk of borrowers, thus reducing the NPLs. For this purpose, we use panel Local Projections of the NPLs-to-output ratios on the ERDF shocks identified in the first stage of the analysis. We also investigate whether the response of NPLs to ERDF shocks may be different in period of tight vs. slack loan supply, and in regions characterized by high vs. low degree of competitiveness, by augmenting the panel Local Projection with a Smooth Transition Function, having as a transition variable either an indicator of the credit supply conditions at the macro-area level or the Regional Competitiveness Index and the three sub-indices pillars developed by the European Commission.

The empirical evidence shows that the ERDF investments have strong short-run effects on the economy of the Italian regions, with multipliers ranging around 2. Furthermore, these investments contribute to the reduction in the financial fragility of the business sector, during the period 2009-2018, mostly characterized by financial crises. This effect is higher in period of credit supply easing, when supply of loans increases and it becomes more likely that riskier borrowers obtain loans from banks. In addition, this effect is also higher in the Italian regions with a lower degree of competitiveness, especially in terms of innovation, thus playing a higher counter-cyclical role when the regions face more real frictions, due to the low level of competitiveness.

The results suggest that policy makers, beyond the use of government guarantees granted to banks, should take into account the beneficial effect of ERDF spending which improve the financial conditions of the business sector especially in regions with a low degree of competitiveness.

Although the results may inform on the potential counter-cyclical role that EU structural funds can play, this analysis has the limitation of not considering issues related to two aspects. First, regional heterogeneity in reacting to shocks. Although we have disentangled the effects between less competitive and more competitive regions, a more thorough investigation of the heterogeneous impact of EU funds should be pursued in future research. Second, as argued throughout the chapter, we have derived a general *"geographic cross-sectional effect"* of ERDF spending, without taking into account the interrelationships between regions and the potential occurrence of spillover effects. This represents another line to extend our analysis further in the future.

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Chapter 3[36]

3 Italian local fiscal multipliers: evidence from proxy-SVAR

3.1 Introduction

In the wake of the financial crisis, European governments have undertaken fiscal consolidation measures to boost economic development and lower the debt-to-GDP ratio. This has often entailed reducing government expenditure and cutting wages for public sector workers, raising taxes, and increasing the retirement age. As a result, economists from various schools have started to question the efficacy of fiscal consolidation measures, saying that austerity will have long-lasting detrimental impacts on current and future production (Fatas and Summers, 2018). It is well established in the fiscal multipliers literature that fiscal stimulus leads to real GDP gains. However, the size of multipliers varies from study to study (Gechert, 2015). For example, Chodorow-Reich (2019) reviews multipliers estimated with regional data and finds that previous studies obtained multipliers with different sizes but mostly positive. Furthermore, since the late 1990s, Italy's post-crisis fiscal austerity policies have gradually reduced the redistributive fiscal flows across regions. As a result, it aggravates rather than reduces interregional imbalances (Giannola et al., 2016; Petraglia et al., 2020). Consequently, Italy was among the countries in Europe hardest hit by the economic crisis (Deleidi, 2022; Caprioli and Momigliano, 2011). In this context, we bring evidence of the positive effects of fiscal policy in Italy, adopting a SVAR approach, exploiting Italian regional data. The literature on applying this approach to Italian regional data is still being determined. We contribute to this literature by employing an identification strategy that has never been applied to the Italian regional government consumption expenditure. In particular, we extend the methodology developed in the previous chapter, to identify nationally-financed government spending shocks. We differentiate from previous studies on Italian regional data, mainly relying on Cholesky identification and sign restrictions (see Deleidi et al., 2021; Destefanis et al., 2022; Lucidi, 2022), using the proxy-SVAR approach. Further, we wonder whether the economic and institutional conditions of the two macro-areas of the country, namely, Centre-North and South, affect the size of the multipliers. Therefore, we extend the VAR model using a dummy variable to disentangle two geographical regimes, thus estimating different multipliers in the Centre-North and the South.

The results show that fiscal policy has positive and long-lasting effects on output, although multipliers are higher in the Centre-North, even though the difference between the multipliers in the two macro-areas is statistically significant in the short-run only. Therefore, our results confirm the Keynesian credo of the

[36]This project has been carried out in collaboration with Marco Maria Matarrese (LUM Giuseppe De Gennaro University, Bari, Italy and Università degli Studi di Bari "Aldo Moro", Bari, Italy). His help is gratefully acknowledged. A short version of this study is published on *Economics Letters* (Matarrese, M. M., and Frangiamore, F. (2023). Italian local fiscal multipliers: Evidence from proxy-SVAR. *Economics Letters, 228*, 111185.). I would like to thank Davide Furceri and Pietro Pizzuto for providing the data on the Italian government expenditure forecast from their paper. I am also grateful to my supervisor Andrea Cipollini and to Konstantin Boss for helpful suggestions.

positive effects of fiscal policy. However, the size of the multipliers is different in the two macro-areas of the country, higher in the more developed area of the Centre-North, confirming previous results pointing at a more effectiveness of fiscal policy in developed countries (see for example Ilzetzki et al., 2013).

3.2 Data

We construct a panel dataset at the NUTS-2 level for Italy, over the period 1995-2019. Data on government spending and output come from the Italian Institute of Statistics (ISTAT). We take the total amount of government consumption expenditure, observed at annual frequency for all the NUTS-2 regions. Data on GDP are also observed at annual frequency for all the NUTS-2 regions. We use the concatenated values with reference year 2015, as measure of real output and real government spending. Table 3.1 shows the average real expenditure and GDP per capita[37], and the average percentage share of expenditure in GDP, over the 1995-2019 period. From this, we can notice that the regions in the South have a smaller GDP per capita with respect to the regions in the North. This reflects the well-known historical North-South divide of Italy, with different levels of economic development in the two macro-areas. This stylized fact is contrasted by the higher percentage share of expenditure in GDP in the South, as can be seen from the last column of Table 3.1.

Table 3.1: Government consumption expenditure (G) and GDP (Y), average 1995-2019, real per-capita.

Region	G	Y	G % of Y
Abruzzo	5440.45	24785.67	21.94
Basilicata	5571.78	20400.81	27.37
Calabria	6504.26	16995.97	38.31
Campania	5244.37	18953.00	27.68
Emilia-Romagna	5068.29	34580.00	14.67
Friuli-Venezia Giulia	5836.28	30265.28	19.30
Lazio	5639.90	34958.41	16.13
Liguria	5657.47	31493.14	17.97
Lombardia	4881.91	37743.71	12.94
Marche	5257.64	26985.35	19.51
Molise	6058.32	21499.34	28.22
Piemonte	5099.25	30406.75	16.79
Provincia Autonoma Bolzano	7993.51	41933.57	19.08
Provincia Autonoma Trento	7829.05	37461.01	20.93
Puglia	5148.29	18089.69	28.46
Sardegna	5929.77	20553.57	28.85
Sicilia	5905.33	18274.74	32.34
Toscana	5217.37	30430.82	17.15
Umbria	5336.96	26870.61	19.96
Valle d'Aosta	9700.19	40074.08	24.38
Veneto	4979.27	31899.19	15.62

We also construct an IV for the output, to be used in the identification strategy (for a more in-depth discussion see section 3.3.3), that consists of an interaction between the share of manufacturing sector value added of the regions with the international oil price. We take data on GVA by sector, at constant 2015 prices, from ISTAT, and we divide the GVA of the manufacturing sector by the total GVA to

[37]Data on population, to compute the variables in per-capita terms, are from ARDECO.

compute the share of manufacturing sector value added. We use the Brent oil price, from the World Bank Commodity Price Data, as measure of the international oil price. Since this is a nominal price in US dollars, we use the CPI of all items for the United States to compute the real oil price. To be consistent with the reference year of the other variables, we take the CPI from the FRED database[38], with the 2015 as base year. Finally, to take into account the "fiscal foresight" issue, we use the IMF forecast of the Italian total government expenditure taken from Colombo et al. (2022). We obtain the series in nominal terms, expressed in Italian lira until 2001 and in euro from 2002 onwards. Therefore, we first divide the series up to 2001 by the exchange rate, set during the negotiations for entry into the Eurozone, which is 1936.27 lire per 1 euro. Then, we use the CPI of all items for Italy, from the AMECO database, with 2015 as base year, to obtain the real expenditure forecast. Since this series is observed at the national level, in an attempt to obtain a measure at the regional level, instead of assuming that the expenditure forecast is homogeneous across the Italian regions (Deleidi et al., 2021), in this research we propose to scale it with the regional share of public expenditure, i.e. we divide each region's expenditure by the total Italian public expenditure and multiply this share by the Italian public expenditure forecast. Therefore, we assume that the regional expenditure forecast is proportional to the national one, with weights given by the regions' share of public expenditure in national public expenditure[39].

Figure 3.1: Average growth rate, relative to GDP lag, of our variables, across the 21 regions.

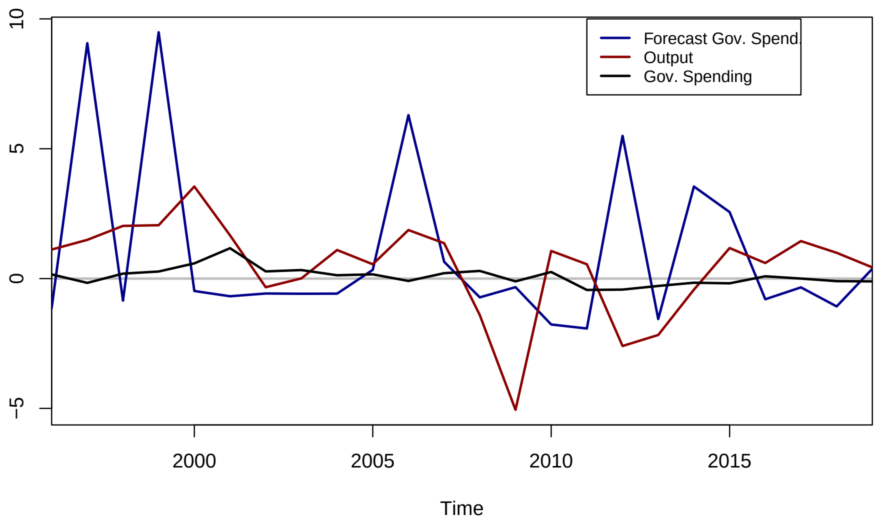


Figure 3.1 shows the evolution of government consumption expenditure, output and government expenditure forecasts. Relevant episodes emerge from this plot: (i) the economic growth of the late 1990s and before the ICT bubble in the early 2000s, the sharp drop in output during the Global Financial Crisis and the Sovereign Debt Crisis, and the recovery after the latter episode; (ii) the need to contain the growth of public spending in order to meet the requirements of the currency union on public debt and deficit at the beginning of our sample, before joining the euro area, and the process of fiscal consolidation

[38] Mantained by the Federal Reserve of St. Louis.

[39] Figure A3.1 in Appendix A3 compares the evolution of the expenditure forecast at the national level with the one calculated for each region.

after the Sovereign Debt Crisis, where the growth rate of public spending was negative from 2011 to 2015; (iii) the fact that forecasters greatly overestimated or underestimated public expenditure growth during periods of high uncertainty, such as the post-1992 lira crisis and the periods surrounding Eurozone entry, the Global Financial Crisis and the Sovereign Debt Crisis (Figures A3.1 and A3.2 in Appendix A3 show, respectively, the evolution of the forecast expenditure at national and regional level and of the output and government spending for each region).

3.3 Emprirical strategy

In order to study the effects of fiscal policy in Italy, and estimate the associated multipliers, we exploit both time series information and regional heterogeneity, by applying a SVAR approach on panel data at the NUTS-2 level. Similar studies adopting this approach on Italian NUTS-2 data are Destefanis et al. (2022), Deleidi et al. (2021), Lucidi (2022) and Zezza and Guarascio (2022). We differentiate from them by applying a different identification strategy, a different way to deal with the fiscal foresight issue, and a different trasformation of the variables that enter the VAR, used in order to compute the multipliers.

3.3.1 Econometric model

We estimate a VAR model (Sims, 1980) on panel data, with two endogenous variables, that are a fiscal variable and a variable for the output, controlling for time and regional fixed effects. The reduced-form model is as follows:

$$y_{i,t} = \alpha_i + \gamma_t + \Gamma y_{i,t-1} + \epsilon_{i,t} \tag{3.1}$$

where $y_{i,t} = [G_{i,t}, Y_{i,t}]$ is the vector of endogenous variables for region i at time t, with $G_{i,t}$ the government expenditure variable and $Y_{i,t}$ the output variable. Since these variables are observed at annual frequency, we estimate the reduced form model with one lag of the endogenous variables, $y_{i,t-1}$. The 2×2 matrix Γ contains the slope coefficients of each VAR equation, whereas α_i and γ_t are regional and time fixed effects, respectively[40]. The vector of the two innovations is $\epsilon_{i,t}$. The innovations are contaminated by all the structural shocks hitting the system. Thus, to make structural analysis, we need to recover the structural form, that we represent as in the B-model (Lütkepohl, 2005), where the innovations are related to the structural shocks by means of a matrix B that contains the on-impact responses of the variables to the structural shocks $u_{i,t}$:

$$\epsilon_{i,t} = B u_{i,t} \tag{3.2}$$

In this paper we are interested in the government spending shock, which involves the identification of the first column of this matrix[41]. Once we identify the structural form, we are interested in the dynamic response of the variables and in the estimation of the multipliers, over a time-span horizon after the occurrence of the shock. Therefore, we retrieve the structural Impulse Response Functions (IRFs), by

[40] We estimate the reduced form applying the OLS estimator on within transformed data to control for time and regional fixed effects.

[41] In line with the second chapter, we follow Nakamura and Steinsson (2014) and Chodorow-Reich (2019) and estimate a pooled panel VAR with regional and time fixed effects in order to obtain the *"geographic cross-sectional fiscal spending multiplier"*. Therefore, the parameters are constrained to be homogeneous across the regions. Also, as already pointed out in the second chapter, the T dimension of the panel dataset may limit the use of other estimation methodology, such as the Pesaran and Smith (1995) mean group estimator.

multiplying the powers of the matrix containing the coefficients of the lags by the matrix B, $IRF_h = \Gamma^h B$, where $h = 0, 1, \ldots, 6$ is the time horizon after the shock. Since they are non-linear functions of the VAR parameters, we implement a bootstrap procedure for the inference. We apply the pairs bootstrap for panel data (Kapetanios, 2008), directly resampling the observation in the time dimension for each region. We perform 1,000 repetitions, and we construct the 68% confidence interval bands by computing the 16-th and 84-th quantile of the simulated IRFs distribution.

3.3.2 Multipliers computation

Different methodology has been used in the VAR literature to compute fiscal multipliers. One approach is the inclusion of the logarithm of the fiscal and output variables into the VAR. In this case the IRFs measures the elasticity of the variables to the government spending shock. To obtain the multipliers, typically the IRF of output is multiplied by a conversion factor given by the sample average of the ratio between GDP and government expenditure (for an application of this approach on Italian regional data see Deleidi et al., 2021). As argued by Ramey and Zubairy (2018), this makes the multipliers potentially dependent on the sample period in which they are estimated, and the authors show that this can bias the estimates. An alternative transformation is the one used by Hall (2009) and Mumtaz and Sunder-Plassman (2020), where the first difference of the variables are divided by the lag of GDP. This is done because it ensures that all the variables are expressed in the same unit, and the multipliers can be estimated by the ratio between the cumulated IRF of GDP and the cumulated IRF of government expenditure to a fiscal policy shock. Therefore, we use the following data transformation: (i) $G_{i,t} = (\tilde{G}_{i,t} - \tilde{G}_{i,t-1})/\tilde{Y}_{i,t-1}$; (ii) $Y_{i,t} = (\tilde{Y}_{i,t} - \tilde{Y}_{i,t-1})/\tilde{Y}_{i,t-1}$. $\tilde{G}_{i,t}$ and $\tilde{Y}_{i,t}$ are, respectively, real government consumption expenditure and real GDP in level. The variables enter the VAR with this transformation, thus we compute the multiplier at horizon h as follows:

$$m_h = \frac{\sum_{j=0}^{h} IRF_j^Y}{\sum_{j=0}^{h} IRF_j^G} \tag{3.3}$$

3.3.3 Identification

To identify the effects of fiscal policy, a bunch of previous studies adopted a recursive scheme on quarterly variables, assuming that government spending is not contemporaneously affected by output, due to time needs in the decision process of the authority and in the implementation of fiscal policy (see for example Blanchard and Perotti, 2002; Giordano et al., 2008 for Italy; and many others). Destefanis et al. (2022) and Deleidi et al. (2021) employed this strategy on Italian regional data at annual frequency. In this chapter we try to avoid this assumption and use a different identification strategy, borrowed from Brueckner et al. (2023). Furthermore, there may be anticipation effects by private agents that invalidate the use of Cholesky identification to impose zero restriction on the feedback of government spending to output, an issue highlighted by the literature about the so-called "fiscal foresight"(Ramey, 2011; Forni and Gambetti, 2010; Leeper et al., 2013). Indeed, the agents react suddenly to fiscal news, before the actual policy takes place, thus they adapt their consumption and investment decisions accordingly, affecting the GDP before the government implements the policy actions. This essentially leads the econometrician to capture an anticipated shock, thus obtaining biased estimates (for a more in-depth and formal discussion see Forni and Gambetti, 2010). We try to address this issue by applying a proxy-SVAR approach (Stock and Watson, 2012; Mertens and Ravn, 2013; Gertler and Karadi, 2015), where the proxy for the fiscal policy shock is an estimated series of government expenditure changes unrelated to the GDP and purged from the expenditure forecast. Therefore, this allows us to obtain a proxy for the structural shock hitting government expenditure, that is exogenous with respect to the other endogenous variable and unanticipated by construction[42]. To obtain such series, we follow Brueckner et al. (2023) and estimate the following panel data model:

$$G_{i,t} = \alpha_i + \gamma_t + \beta Y_{i,t} + \delta FG_{i,t|t-1} + \rho X_{i,t-1} + r_{i,t} \tag{3.4}$$

where $G_{i,t}$ and $Y_{i,t}$ are government expenditure and output, transformed as showed in section 3.3.2; α_i and γ_t are, respectively, regional and time fixed effects; $X_{i,t-1}$ is a vector of control variables that includes a lag of $G_{i,t}$ and $Y_{i,t}$, to control for past states of the economy and fiscal policy, that can have effect on future implementation of fiscal policy. $FG_{i,t|t-1} = [(\tilde{FG}_{i,t} - \tilde{FG}_{i,t-1})|t-1]/\tilde{Y}_{i,t-1}$ is the forecast of government spending made by the agents for year t, based on information available up to year t-1, which is transformed in the same way as the other variables, namely, the first difference is divided by the lag of GDP. Therefore, this variable represents the agents' forecast of changes in government expenditure, conditional on information available up to the previous year, so we can rule out any endogeneity problems arising from the inclusion of this variable in equation (3.4). We take the estimated residual of this model as proxy for the government expenditure shock. In the way in which model (3.4) is specified, the residuals are purged from current and past macroeconomic conditions, here captured by the presence of contemporaneous and lagged GDP, and from the predictable component, since we also include the forecast of the expenditure among the regressors. However, $Y_{i,t}$ is endogenous in this model, because of the reverse causality, namely, output is influenced by government expenditure. Thus, we follow Brueckner et al. (2023) and we use the IV approach to estimate equation (3.4). The IV is constructed as in Brueckner et al. (2023), exploiting variations in the international oil prices[43]. In particular, we construct the IV as interaction between the growth rate of the international oil prices and the regional share of manufacturing value added. The use of the international oil price is due to its important role in shaping the output dynamics of every country and region, especially those that use oil most intensively as a source of energy and raw material. For this reason we compute the share of manufacturing GVA of each region in each year, and multiply this share by the growth rate of the international oil prices. As stated above, the endogeneity is due to the reaction of GDP to fiscal policy, thus our instrument need to be exogenous to government spending. Since the time variation of our IV is driven by changes in the international oil prices, we can assume that it can hardly respond to fiscal policy conducted at the Italian regional level. However, a source of endogeneity of this instrument can be represented by the share of the manufacturing sector, that

[42]This approach is the same used in the second chapter, but it is extended by including the forecast of government spending which is available for the nationally-financed government expenditure but not for the EU structural funds.

[43]Other studies that estimate similar models are Alesina et al. (2008) and Fatás and Mihov (2013). In particular, Fatás and Mihov (2013) exploit variations in the international oil prices to instrument output.

can be affected by government expenditure. We estimate a panel data model, having the IV as dependent variable and government spending as explanatory variable. We also control for time and regional fixed effects, as we do throughout the paper. The results in Table 3.2 show that the IV does not depend on government expenditure. This result is robust to any kind of violation of the classical assumptions on the residuals. Indeed, we apply White (1980), White (1984) heteroskedasticity corrections, Arellano (1987) heteroskedasticity and serial correlation correction and Driscoll and Kraay (1998) heteroskedasticity, serial correlation and cross-sectional dependency corrections, to compute standard errors. The coefficient of government spending is close to zero and not statistically significant.

Table 3.2: Results of the regression of our IV on government spending ($G_{i,t}$). The p-values are computed using different estimators of the standard errors, which are listed in the first column.

Estimator std. errors	variable	coefficient	p-value
Classic	$G_{i,t}$	-0.21	0.33
White (1980)	$G_{i,t}$	-0.21	0.34
White (1984)	$G_{i,t}$	-0.21	0.30
Arellano (1987)	$G_{i,t}$	-0.21	0.24
Driscoll and Kraay (1998)	$G_{i,t}$	-0.21	0.53

Of course, this is not a test on the exogeneity, and it is performed just to reassure that in our sample government spending does not predict our IV. In section 3.4.3, we conduct a robustness exercise to further endorse our analysis by constructing a Bartik-type instrument (Bartik, 1991), where the share of manufacturing value added is fixed at the 1994 year. In this way, fiscal policy conducted by the Italian regions from 1995 to 2019 cannot influence the GVA share of the manufacturing sector in 1994.

We test the relevance of our IV, taking the estimated F-statistic from the first stage regression. We use some robust estimator proposed in the literature, in particular, the Arellano (1987) estimator, that takes into account serial correlation and heteroskedasticity, and also, we compute the Kleibergen-Paap F-statistic, as suggested by Andrews et al. (2019). Table 3.3 shows the results of the first stage regression. The F-statistic computed using the Arellano (1987) estimator is above the threshold of 10 suggested by Stock et al. (2002), and the Kleibergen-Paap F-statistic is equal to 15.15, above the threshold of Stock and Yogo (2005) for a 15% maximal IV size. This ensures that the relevance condition is satisfied.

Table 3.3: First stage results of equation (3.4).

Variables	coefficients	std. errors
$IV_{i,t}$	0.13***	0.028
$FG_{i,t}$	0.20***	0.05
$G_{i,t-1}$	0.29	0.15
$Y_{i,t-1}$	-0.08*	0.03
F-statistic	20.04	
Kleibergen-Paap statistic	15.15	

Note: standard errors robust to heteroskedasticity and serial correlation (Arellano ,1987). *** and * indicate statistical significance at the 0.1% and 5% level, respectively.

The results obtained from the estimation of equation (3.4) are presented in Table 3.4. It shows that government spending was counter-cyclical, since the coefficient of output is negative, even though it is not statistically significant[44]. The forecast of government spending has a significant positive impact on

[44]As also argued in the second chapter, this is not a test on the exogeneity of government spending to output within a

Table 3.4: Results of IV estimation of equation (3.4).

Variables	coefficients	std. errors	t-value	p-value
$Y_{i,t}$	-0.0060584	0.0079012	-0.7668	0.443635
$FG_{i,t}$	0.1474763	0.0322962	4.5664	6.466e-06***
$Y_{i,t-1}$	0.0238359	0.0131021	1.8193	0.069559#
$G_{i,t-1}$	0.0843501	0.0280101	3.0114	0.002751**

Note: Note: standard errors robust to heteroskedasticity and serial correlation (Arellano ,1987). ***, ** and # indicate statistical significance at the 0.1% and 1% level and 10% levles, respectively.

actual government spending. Finally, the effects of past states of the regional economies are positive on future regional government expenditure, which, however, are statistically significant at the 10% level only, and lags of government spending significantly affects the evolution of future regional government expenditure.

Finally, other comments are in order. There could be different confounding factors in the estimation of the effects of fiscal policy on output. For example, monetary policy may be driven by the evolution of fiscal policy and it also affects output, so monetary policy shock is correlated with both output and government expenditure. The omission of a monetary policy variable in the vector of endogenous variables may bias the estimates. However, the inclusion of time fixed effects can somehow help in dealing with this issue. Indeed, monetary policy is common across the Italian regions, and its effects may be captured by the time fixed effects.

Moreover, the inclusion of time fixed effects allows to control for other aggregate shocks, like global shocks. Furthermore, the regional fixed effects control for characteristics of the Italian regions that are stable over time, like structural differences, that can be relevant, especially between the Northern and the Southern ones. Therefore, by construction and by the assumption discussed above, our proxy can fulfill the characteristics of a structural shock indicated in Ramey (2016): (i) it should be exogenous to the contemporaneous and lagged endogenous variables, as model (3.4) include current and past values of the endogenous variables; (ii) it should be unrelated to the structural shock hitting GDP, as by application of the IV estimator in equation (4); (iii) it should be unanticipated, as it is purged from government expenditure forecast. For these reasons we can assume that our proxy is exogenous. We test the relevance of our proxy, following Gertler and Karadi (2015) and Caldara and Kamps (2017), namely, we compute the F-statistic of the regression of the policy innovation, $\hat{u}^p_{i,t}$, on the proxy, $\hat{r}_{i,t}$, and we find that it is equal to 1058. It is well above the threshold of 10 suggested by Stock et al. (2022), thus it ensures that the relevance condition is satisfied. Then, this proxy is used in the application of the proxy-SVAR approach. Especially, as in the second chapter, we follow the model notation of Gertel and Karadi (2015), thus, we denote with $u^p_{i,t}$ the policy shock hitting government expenditure, and with $u^q_{i,t}$ the output shock. As we said, the estimated residuals from equation (3.4), $\hat{r}_{i,t}$, serve as a proxy for the government expenditure shock $u^p_{i,t}$. The application of the proxy-SVAR involves the following steps:

- estimate the reduced-form VAR (1) and the innovations, $\hat{\epsilon}^p_{i,t}$ for the equation of government expenditure, and $\hat{\epsilon}^q_{i,t}$for the equation of GDP;
- regress the policy innovation $\hat{\epsilon}^p_{i,t}$ on the proxy $\hat{r}_{i,t}$ and compute the fitted values (first stage), that,

year. Therefore, this can not be used as evidence in favour of the application of the Cholesky restrictions.

as discussed above, should represents the variations in government spending that are exogenous to the other shock, to current and lagged endogenous variables and unanticipated. Let us indicate these fitted values with $\hat{u}^p_{i,t}$;

- estimate the response of output to the government spending shock, by regressing the output innovations $\hat{\epsilon}^q_{i,t}$ on the fitted values $\hat{u}^p_{i,t}$:

$$\hat{\epsilon}^q_{i,t} = \frac{B_{2,1}}{B_{1,1}} \hat{u}^p_{i,t} + \phi_{i,t} \tag{3.5}$$

- up to the normalization of the response of government spending to its own shock $B_{1,1} = 1$ (first element of first column of matrix B), we identify the first column of matrix B, that is $B_{\bullet,1} = [1, B_{2,1}]'$.

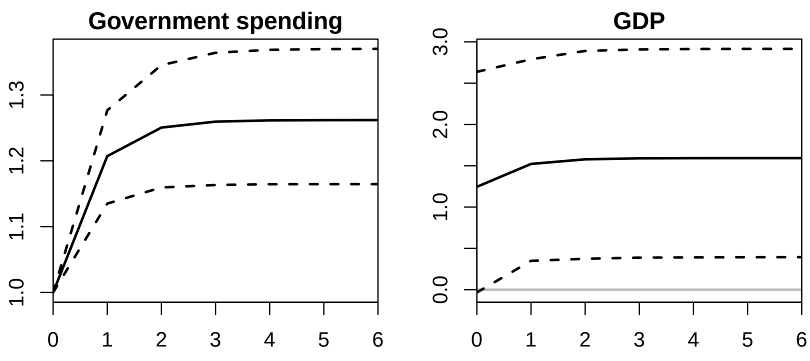
Now, we can estimate the structural IRFs and the multipliers, as discussed in section 3.3.1 and 3.3.2. All the steps of the analysis, including the estimation of equation (3.4), are included in the bootstrap.

3.4 Results

3.4.1 Baseline results

Figure 3.2 shows the cumulative IRFs of government expenditure and output to a government spending shock over a 6 years time span. From this picture we can see that a government spending shock has a significant and long-lasting effect on the level of government expenditure and GDP. Output raises significantly one year after the shock, and then the effects are persistent for all the subsequent years. The on-impact response is not significant. These results are in line with previous studies on fiscal policy for Italy, finding that government expenditure significantly stimulates the economy (see Giordano et al., 2007; Deleidi et al., 2021; Deleidi, 2022; Destefanis et al., 2022).

Figure 3.2: Cumulative Impulse Response Functions (solid lines) and 68% confidence interval bands (dashed lines).



The cumulative multipliers, estimated using equation (3.3), are in Table 3.5. They reflect the results obtained from the IRFs analysis, by construction. They are all positive and statistically significant from horizon 1 onwards. A one euro increase in government consumption expenditure generates a cumulative output gain of 1.26 euros after 6 years. Deleidi et al. (2021) estimate a similar model using a different identification strategy. They assume orthogonality of government consumption to output shocks, controlling for the expenditure forecast. They find a cumulative multipliers for government consumption of 1.84, six year after the shock. Our estimates are lower than theirs, and the difference can be due to the identification approach and the transformation of the data adopted to compute the multipliers.

Table 3.5: Cumulative multipliers.

horizon	0	1	2	3	4	5	6
multipliers	1.245287	**1.261119**	**1.262009**	**1.262386**	**1.262441**	**1.262456**	**1.262458**

Note: Statistical significant multipliers in bold, according to the significance of IRFs.

3.4.2 Geographical heterogeneity of Italian regional multipliers

A well-known stylised historical fact is the North-South divide of Italy. Since the unification of the country, the two macro-areas have been characterised by different social and economic conditions. The gap between the two macro-areas still persists and sees the South lagging behind in terms of economic performance and institutional quality (Fina et al., 2021; Di Liberto and Sideri, 2015). It is important to study how this geographical heterogeneity may affect the effectiveness of fiscal policy in Italy (Deleidi et al., 2021). The worse institutional conditions and worse economic performance in the South may lead to relatively lower efficiency in the use of public sources, resulting in lower fiscal multipliers in this less developed area (Petraglia et al., 2020; Zezza and Guarascio, 2022). This result seems to be prevalent in the Italian case, as showed by some empirical analysis on Italian regional data (see Deleidi et al.,2021; Lucidi, 2021). Also, studies involving country-level data, find that in less developed countries fiscal multipliers are lower than the ones estimated for the advanced economies (Ilzetzki et al., 2013). This results also hold for the European regions, as showed in Gabriel et al., (2023), that finds lower multipliers in peripheral EU regions. Therefore, we explore whether fiscal multipliers estimated in section 3.4.1, differ between Centre-North and South. To do so, we extend the model presented in section 3.3.1 and 3.3.3, to allow the IRFs and multipliers to vary across the two macro-areas.

Extension of the model. We group the regions according to the NUTS-1 classification. The Central-Northern area include regions belonging to the North-East, North-West and Centre. They are the following: Piemonte, Valle d'Aosta, Liguria, Lombardia, Emilia-Romagna, Friuli-Venezia Giulia, Veneto, Provincia Autonoma di Bolzano, Provincia Autonoma di Trento, Veneto, Lazio, Marche, Toscana and Umbria. The Southern area merges regions in the South and the Islands, that are Abruzzo, Basilicata, Calabria, Campania, Molise, Puglia, Sardegna and Sicilia. According to these groups, we construct a dummy variable taking 1 if the region belongs to the South group, and 0 otherwise. We, then, interact this dummy variable with the variables in the right-hand-side of the VAR equations, i.e., we estimate a dummy-augmented VAR model, whose reduced-form is as follows:

$$y_{i,t} = \alpha_i + \gamma_t + \Gamma_S[D_i \times y_{i,t-1}] + \Gamma_N[(1 - D_i) \times y_{i,t-1}] + \epsilon_{i,t} \tag{3.6}$$

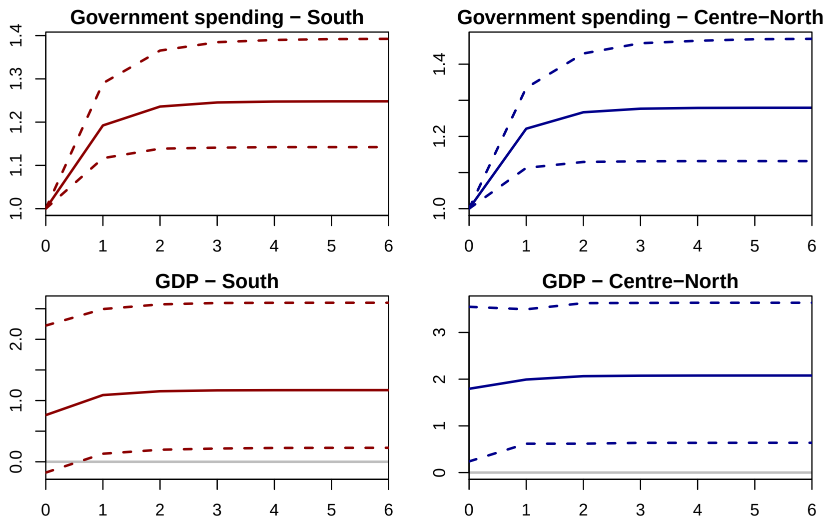
where D_i is the dummy variable, thus Γ_S is the matrix of VAR coefficients of the South, whereas Γ_N is the matrix of coefficients of the Centre-North. We allow for different impact coefficients, by separating the shock $\hat{u}^p_{i,t}$, obtained in the second step of the proxy-SVAR, between the two areas, using its interaction with the dummy variable, in equation (3.5) that now is as follows:

$$\hat{\epsilon}^q_{i,t} = \frac{B^S_{2,1}}{B^S_{1,1}}[D_i \times \hat{u}^p_{i,t}] + \frac{B^N_{2,1}}{B^N_{1,1}}[(1 - D_i) \times \hat{u}^p_{i,t}] + \phi_{i,t} \tag{3.7}$$

Again, up to normalization to 1 of $B^S_{1,1}$ and $B^N_{1,1}$, namely the response of government expenditure to its own shock, respectively in the South and Centre-North, we identify the column of interest of matrix B. Therefore, we have $B^S_{\bullet,1} = [1, B^S_{2,1}]'$, that is the vector containing the on impact responses of our endogenous variables to a spending shock in the South, whereas $B^N_{\bullet,1} = [1, B^N_{2,1}]'$ in the centre-North. Consequently, $IRF^S_h = \Gamma^h_S B^S_{\bullet,1}$ and $IRF^N_h = \Gamma^h_N B^N_{\bullet,1}$ are, respectively the IRFs for South and Centre-North, and thus the multipliers for the two macro-areas are estimated using equation (3.3), applied to the latter IRFs, distinguished by area. This method differs from the one applied by Deleidi et al. (2021) and Zezza and Guarascio (2022), where the data are divided into two sets, one for each macro area, and the model is estimated for each of these two sub-samples. The two approaches should give similar results, but the way we proceed allows us to conduct tests on the significance of the difference of the multipliers between the Centre-North and the South, computing this difference within the bootstrap. In this way we obtain a simulated distribution for this difference and calculate the 68% confidence interval bands.

Results on the geographical heterogeneity of the multipliers. Figure 3.3 shows the IRFs of government expenditure and output to fiscal shocks in the South (in red), and in the Centre-North (in blue).

Figure 3.3: Cumulative Impulse Response Functions by macro-area (solid lines) and corresponding 68% confidence interval bands (dashed lines).



A positive fiscal shock leads to persistent effects on government expenditure and output, in both areas of the country. However, multipliers differ across the two macro-regions. Table 3.6 shows the estimated multipliers (in bold we again indicate the statistical significant multipliers according to the significance of the IRFs). In line with Deleidi et al. (2021) and Lucidi (2022), we find higher multipliers in the

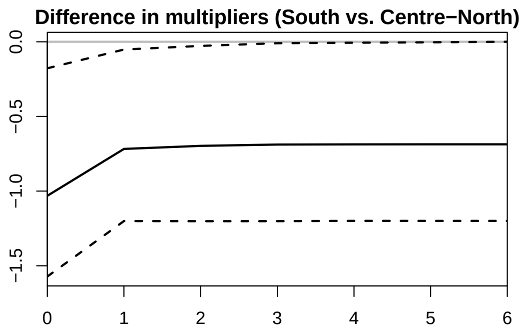
Centre-North. A 1 euro increase in government spending determines a cumulative gain of 1.62 euros, after 6 years, in the Centre-North, against 0.94 euros in the South. Again, our estimates are lower than the ones obtained by Deleidi et al. (2021), both for the Centre-North and the South. Furthermore, our estimates for the Centre-North are higher than the estimates obtained by Lucidi (2022), whereas for the South are lower than the ones estimated by the same paper. Figure 3.4 shows the point estimate of the difference between the multipliers in the two macro-areas, along with the 68% confidence interval bands. It shows that this difference is significant in the short-run only.

Table 3.6: Heterogeneous (South vs. Centre-North) cumulative multipliers.

year	0	1	2	3	4	5	6
South	0.762498	**0.913809**	**0.932188**	**0.936515**	**0.937427**	**0.937628**	**0.937672**
Centre-North	**1.794023**	**1.631329**	**1.629237**	**1.624777**	**1.624490**	**1.624324**	**1.624306**

Note: Statistical significant multipliers in bold, according to the significance of IRFs

Figure 3.4: Difference between multipliers in the South and multipliers in the Centre-North (solid line) with 68% confidence interval bands (dashed lines).



3.4.3 Robustness check

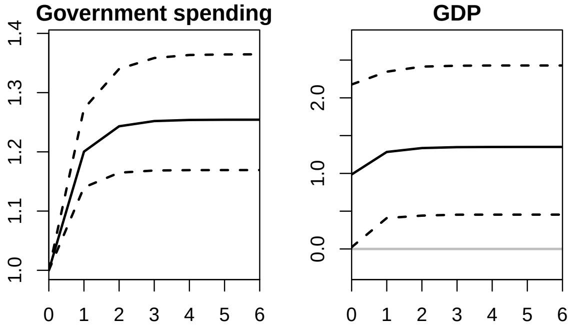
We perform a robustness exercise, changing the way in which we construct the IV for output in equation (3.4). As stated in section 3.3.3, the regional shares of manufacturing value added can be influenced by regional fiscal policy. To overcome this issue, we compute a static share in a pre-sample in 1994. This ensures that fiscal policy conducted from 1995-2019 (our sample) does not affect the value-added composition of each region in 1994. The Kleibergen-Paap rk Wald F-statistic of this new IV equals 21.57, revealing that it satisfies the relevance condition. The F-test for the relevance of the proxy, computed as in Gertler and Karadi (2015), is equal to 1176. Figure 3.5 and Table 3.7 show that the results are qualitatively the same, and the multipliers are slightly lower but still positive and above the unity.

Table 3.7: Cumulative multipliers estimated using the IV constructed in the robustness check (see section 3.4.3).

horizon	0	1	2	3	4	5	6
multipliers	0.985777	**1.069426**	**1.07416**	**1.076165**	**1.076458**	**1.076533**	**1.076547**

Note: Statistical significant multipliers in bold, according to the significance of IRFs.

Figure 3.5: Cumulative Impulse Response Functions (solid lines) and 68% confidence interval bands (dashed lines) estimated using the IV constructed in the robustness check (see section 3.4.3).



3.5 Conclusions

In this study, we bring evidence on regional fiscal multipliers in Italy. We exploit a panel of 21 Italian NUTS-2 regions from 1995 to 2019. We apply a proxy-SVAR approach, instrumenting the government spending shock with a series of spending changes unrelated to GDP and purged from predictable components. This series is estimated in the first stage, where we deal with the endogeneity of output using the IV estimator, where the instrument for the output is constructed as interaction between the regional share of manufacturing value added and changes in the international oil prices. Our results show that public spending boosts the economy, by stimulating the output and generating a cumulative multipliers of about 1.26 after 6 years. Furthermore, we explore the presence of geographical heterogeneity of multipliers, that is, we estimate multipliers for two macro-areas, the Centre-North and the South of Italy. To do so, we extend the model and apply a dummy-augmented VAR, where the dummy variable takes one if a region belongs to the South and zero otherwise. We find that multipliers are higher in the Centre-North. We test the significance of this difference by including its computation in the bootstrap. This procedure yields a statistically significant difference only in the short run, i.e., up to 2 years after the shock. These results may have important policy implications. First, the Italian government should consider fiscal policy as a tool to stimulate the economy of the Italian regions, given evidence of positive government spending multipliers. Moreover, although we do not study the issue empirically, Chian Koh (2017) suggests that multipliers are lower in poorer countries because of lower efficiency in public expenditure management. Lucidi (2022) argues that this idea can be extended to the North-South divide of Italy. Although providing empirical evidence on why this difference exists is different from the goal of this study, this could be a stimulus for further research on this topic.

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Appendix to chapter 3: additional figures

Figure A3.1: Percent growth rate of forecast expenditure relative to GDP, at country level (red solid lines), and at regional level (black solid lines) computed by the approach proposed in this paper, scaling the national forecast expenditure with the regional share of government expenditure. The horizontal gray solid lines indicate the line at zero.

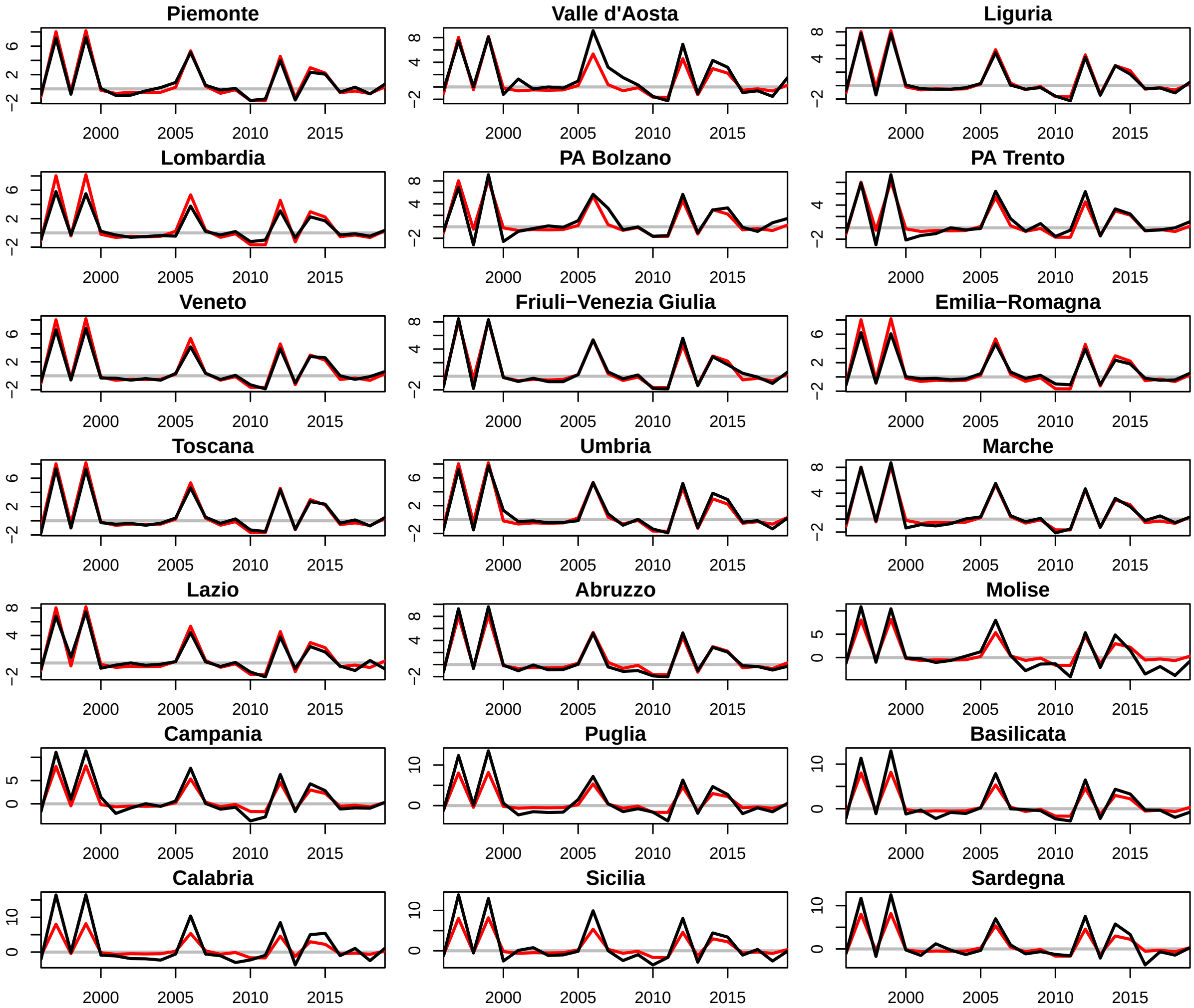
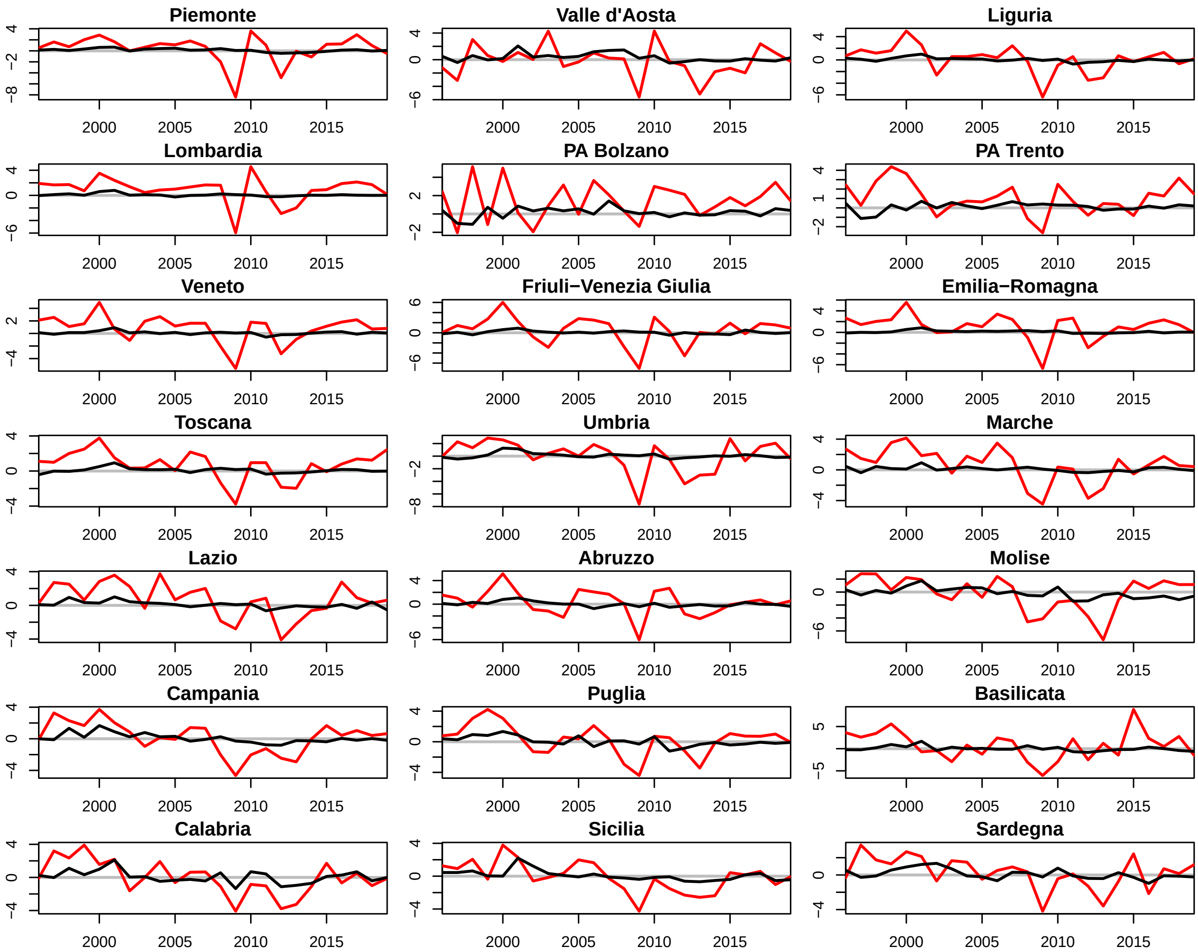


Figure A3.2: Percent growth rate of GDP (red solid lines) and percent growth rate of government spending relative to lags of GDP (black solid lines). The horizontal gray solid lines indicate the line at zero.



Chapter 4

4 Fiscal policy effects on housing prices and credit conditions: evidence from Italian NUTS-2 data

4.1 Introduction

The focus of this study is the analysis of the impact of fiscal policy on house prices in Italy using regional data, available at NUTS-2 level. To our knowledge, the investigation of fiscal policy effects on asset prices (mainly stock and house prices) has been carried at country-level only (see, e.g., Afonso and Sousa 2011, 2012, 2013; Agnello and Sousa 2013; Khan and Reza 2013, 2017; Gupta et al. 2014; Aye et al. 2014; and Ruiz and Vargas-Silva 2016) and the study of local fiscal policy effects has mainly focused on the estimation of fiscal multipliers for output (see Chodorow-Reich 2019, for an extensive review).

The literature on the effects of fiscal policy on house prices has pointed at a direct and an indirect channel of transmission. As for the former, governments can provide subsidies for the purchase of new houses, stimulating demand for houses, pushing up their prices (given an inelastic housing supply). Moreover, on the tax side, as suggested by Afonso and Sousa (2012, 2013) and Agnello and Sousa (2013), governments can impose VAT on the purchase of new houses, taxes on the so-called notional rental value, i.e. the value of owner-occupied houses that add up to taxable income. As for the indirect channel, an expansionary fiscal policy can increase disposable income, hence raising the demand of houses and their prices (Agnello and Sousa 2013; Gupta et al. 2014): this is the income channel. In addition, another indirect channel work through the influence on credit market conditions. The effects of government spending on long-term interest rates are transmitted to the interest rate on mortgages, and thus to the demand for houses. In particular, an expansionary fiscal policy, triggering a confidence crisis on public debt sustainability, has an impact on the sovereign risk premium and can negatively affect the access to mortgages. On the other hand, the studies of Auerbach et al. (2020), based on municipal data, and of Murphy and Walsh (2022), show that, in contrast to the standard macroeconomic literature, there is evidence, in the US, of a positive effects of fiscal policy on credit markets (see also Cipollini and Frangiamore 2023 for an evidence based on Italian NUTS-3 data). Finally, as suggested by Ruiz and Vargas-Silva (2016), government investment in infrastructure, such as roads, public transport, schools, universities, parks, can make houses in a given area more attractive (Ruiz and Vargas-Silva 2016). Furthermore, as suggested by Jappelli and Pistaferri (2007), measures such as the tax deductibility of mortgage interest payment can have a positive impact on the demand for mortgages.

We contribute to the literature on the impact of local fiscal policies by investigating their average effect on house prices, controlling for unobserved heterogeneity through the estimation of a fixed effects Panel VAR. We focus on economy wide style shocks to fiscal policy (Egan and Bergin, 2023), either to the spending and revenue side, and analyse their effects on the housing market and credit market conditions for households[45]. We also analyze these fiscal policy effects conditioning on group of regions differing in terms of the level of economic and banking sector development.

Another contribution is to take into account the interplay between household credit market conditions and housing markets at regional level. More specifically, the causal linkage between credit market condition and house prices is not uni-directional (e.g. via a credit channel transmission mechanism described above) but bi-directional, given the role played by the housing collateral in granting credit to households. In particular, changes in house prices impact the stock of wealth, and houses can be used as collateral for loans, so their value has an influence on the probability of households being subject to credit rationing (Khan and Reza 2013, 2017). We study the interplay between house prices and credit market conditions, by treating them as endogenous variables and by examining a different model specification according to the category of loans: (i) mortgages; (ii) loans for durable goods; (iii) consumer credit.

Finally, we also contribute to the literature on local fiscal policy effects on house prices and credit market conditions disentangling the role played by shocks to government spending and those to tax revenues.

The panel VAR model used is fitted to NUTS-2 Italian data from 2004 to 2019 for government spending and tax revenues, output, proxies of credit market conditions such as loans and interest rates, and house prices. The empirical model is estimated through Bayesian techniques and the identification of the shock to government spending and tax revenues (e.g. the exogenous component of fiscal policy) is obtained through sign restrictions. While the results at country based level, provided by previous studies, are mixed and depend on the country under scrutiny, the identification approach used and the prevailing indirect channel through which fiscal policy influences real estate markets, the empirical evidence of this study shows that an increase in government spending and a tax cut lead to an increase in house prices and to a relaxation of the credit market conditions, fostering the provision of loans. However, we also show that these results depend on some characteristics of the regional economies and on the period under scrutiny. In particular, we provide evidence of more effectiveness of fiscal policy in the regions with a higher level of economic development and with a more developed banking sector. Then, we observe that during the Global Financial Crisis and the Sovereign Debt Crisis, fiscal policy played a stabilizing role not only for output, but also for credit market conditions and house prices.

The remainder of the chapter is organized as follows. Section 4.2 is about the data used in this analysis. Section 4.3 describes the econometric strategy employed and the identification of the shocks, with additional details in Appendix A4 and B4. Section 4.4 shows the baseline results, whereas other results obtained from robustness check exercises are in Appendix D4. Section 4.5 provides evidence on the heterogeneity of the results due to regional characteristics and the period under investigation. Finally, section 4.6 concludes.

[45]Egan and Bergin (2023) define a general fiscal shock, which is not targeted at specific objectives, e.g. to stimulate the housing market, as an economy wide style fiscal shock, which is basically the fiscal shock studied in the macroeconomic literature and analysed by the studies reviewed in section 4.1. However, they also study the effects of a measure, proposed by the Irish government, aimed at stimulating the supply of houses. Although this is not the focus of this chapter, we acknowledge that these two types of shocks may have different impacts on the housing market.

4.2 Data

We construct a panel dataset, at annual frequency, composed by 21 NUTS-2 Italian regions over the period 2004-2019. The variables that we employ are taken from different sources.

Data on regional government expenditure and regional government revenues are from the government agency of territorial cohesion ("Conti pubblici territoriali"). We measure government expenditure as the sum of current and capital expenditure, minus the expenditure on interest rates and current and capital transfers. Government revenues are computed as the sum between direct and indirect tax revenues.

We take the GDP, as measure for the economic activity, from the Italian Institute of Statistics (ISTAT).

Data on house prices are from the real estate market observatory of the Italian revenue agency (Agenzia delle Entrate - Osservatorio del Mercato Immobiliare OMI). This database provides detailed information on the minimum and maximum value of housing at the municipal level and on a semi-annual basis in Italy, for different types of housing and for different structural conditions of housing. The residential house price in a given region is obtained by first computing, for each municipality, the mean of five different typologies of houses, varying from the most affordable to the luxury ones and then by averaging across the municipalities in a specific NUTS-2 region[46]. Moreover, we obtain a series of house prices at semi-annual frequency, where the value is the minimum and maximum price per square meter. We take the average between minimum and maximum as measure of residential house prices and we also average over the two semesters of each year to obtain a measure at annual frequency[47].

Data on the credit market variables are taken from the Bank of Italy ("Base Dati Statistica"). We consider the loans granted to households for different purposes, namely, buying houses, buying loans for durable goods and buying consumption goods. In addition, we take data on the long-term interest rate, charged on household loans. These data are at quarterly frequency, then we average the four quarter of each year to obtain a measure at annual frequency[48].

All these variables are at current prices, thus we use the GDP deflator at the national level to get real measures. Table 4.1 shows the average over the 2004-2019 period.

[46] More specifically, the different typologies of houses considered from the OMI dataset are: (i) Abitazioni civili; (ii) Abitazioni di tipo economico; (iii) Abitazioni signorili; (iv) Abitazioni tipiche dei luoghi; (v) Ville e villini

[47] It is important to observe that the empirical findings are robust to considering separately the minimum and the maximum price (see Appendix D4 on robustness checks).

[48] As for the loans, we use the table TDB10420 and TFR10420, that provide data on mortgages and loans for durable goods, available at the NUTS-2 level. Data on consumer loans are from the table TFR10254, which are only available at the regional level, thus they are not available for the two autonomous provinces of the Trentino-Alto Adige region. Therefore, when we include the consumer credit into the model, we work on 20 Italian regions. Data on interest rates are from the tables TDB30880, TRI30880 and TRI30881. Even in this case the data are not available for the two autonomous provinces of the Trentino-Alto Adige region. Therefore, we repeat the regional data on interest rates for the two autonomous provinces when we work with the panel of 21 NUTS-2 regions, thus giving to these two autonomous provinces the prevailing interest rates in their region.

Table 4.1: Summary of regional variables, average over the 2004-2019 period.

NUTS-2 region	Gov. Exp.	Gov. Rev.	GDP	House prices	Mortgages	Durables	Consumer	Int. rate
Abruzzo	6692.59	7391.73	24632.08	848.66	3060.59	493.55	1082.14	3.16
Basilicata	6334.53	5850.50	21130.15	638.16	1458.63	322.62	920.12	3.06
Calabria	6558.27	5254.58	17448.22	632.24	1549.56	498.98	1151.14	3.23
Campania	5630.46	5783.74	18875.48	1120.39	2474.39	395.21	1050.17	3.18
Emilia-Romagna	6172.39	11010.49	34675.53	1487.26	5967.00	409.51	956.57	2.97
Friuli-VG	8149.32	9483.87	30493.33	939.99	4958.31	375.05	943.18	3.09
Lazio	9804.46	10934.23	35370.90	1649.34	6338.98	535.43	1303.69	3.13
Liguria	6989.56	10154.41	31310.41	2187.89	5474.82	381.21	932.22	3.00
Lombardia	5779.84	12087.92	38015.99	1360.32	6839.11	455.21	1002.13	2.89
Marche	6106.01	8479.03	27206.31	1183.65	4045.79	403.33	937.83	2.99
Molise	7199.13	6443.38	21506.73	679.73	2030.38	308.34	954.93	3.17
Piemonte	6033.89	9930.53	30497.84	1062.49	4931.72	464.93	1113.85	3.01
PA Bolzano	10189.34	11701.75	43025.03	2463.47	3592.22	298.25	630.33	2.85
PA Trento	9585.08	10138.19	37408.76	1988.42	4419.08	307.44	630.33	2.85
Puglia	5581.59	5846.30	18214.08	936.95	2868.43	402.18	1010.64	3.15
Sardegna	7236.90	6400.50	20863.59	1031.63	3179.60	638.92	1355.13	3.27
Sicilia	6332.42	5615.39	18264.60	824.20	2506.57	442.29	1232.31	3.33
Toscana	6264.07	9566.46	30641.90	1872.48	5445.05	507.45	1054.95	2.83
Umbria	6610.41	8237.53	26236.27	1047.88	3413.20	491.92	1135.47	3.11
Valle d'Aosta	12222.27	12341.18	38860.24	2077.29	3742.13	474.01	1127.09	2.63
Veneto	5661.31	9701.27	32083.45	1254.65	5226.66	389.46	870.82	2.98

Note: Durables stands for loans for durable goods, whereas Consumer stands for consumer credit. Government expenditure, government revenue, GDP and loans are in real per capita terms. House prices are in real euros per square meter. The last column shows the long-term real interest rate applied to households loans.

We also employ variables at the country level in the robustness checks, to control for aggregate shocks, which are thus treated as exogenous variables. These are the long-term sovereign bond yield, the short-term interest rate, and the GDP price deflator from FRED to control for the monetary policy stance, and the crude oil price Brent also from FRED, to control for global shocks.

4.3 Empirical strategy

The analysis of the effects of the identified government spending and tax shocks on house prices and credit market conditions is achieved through impulse response analysis based on the estimation of a Panel VAR. The model and the estimation methodology are described in the following sub-sections.

4.3.1 Reduced form model and estimation.

We estimate a Panel VAR(1) with time and regional fixed effects. The reduced form specification is represented as follows:

$$Y_{i,t} = \Phi Y_{i,t-1} + CW_{i,t} + u_{i,t} \tag{4.1}$$

$Y_{i,t}$ is the vector of endogenous variables containing government expenditure, government revenues, output, house prices, households loans and long-term interest rates on loans. We estimate three separate models for each category of credit: mortgages, loans for durable goods and consumer credit.

The model coefficients are described as follows. Matrix Φ is the one containing the VAR slope coefficients; $W_{i,t}$ and C are the vector of exogenous variables and the associated matrix of coefficients, respectively. In the baseline model, the exogenous variables in the panel model specification are regional dummies and time dummies. We also include, among the exogenous variables, a linear and a quadratic time trend to be consistent with the empirical literature working on log-levels variables (Blanchard and Perotti 2002; Khan and Reza 2017). We also conduct some robustness exercises, such as replacing the time dummies with country-level variables or using the within transformation of the endogenous variables to estimate the fixed effects model instead of including the regional and time dummies (see Appendix D4). The vector of innovations is represented by $u_{i,t}$, which is assumed to be normally distributed with 0 mean and constant covariance matrix, namely, $u_{i,t} \sim N(0, \Sigma)$.

The Panel VAR in equation (4.1), estimated through Bayesian techniques, is fitted to variables in log-levels except the interest rate (entering in levels). In particular, we use the Normal-Wishart prior implemented through dummy observations or artificial data (Banbura et al. 2007). The posterior distributions for the reduced form and structural form coefficients are obtained by employing a Gibbs sampling algorithm. We perform 5,000 iterations, discarding the first 4,000 and then making inference with the last 1,000 iterations. We provide details of the estimation and proof of convergence of the algorithm in Appendix A4. The use of transformed data in log-levels is accommodated by incorporating the prior information into the model that macroeconomic variables are $I(1)$ (random walk), in the spirit of the Minnesota prior (Litterman 1986)[49].

Since our panel is small and the sample period is very short, equation (4.1) is estimated by pooling the data (see Canova and Ciccarelli 2013 for a survey on panel VAR estimation), which means that we assume homogeneity of coefficients across the regions[50]. This assumption will be partially relaxed, by considering different sub-samples. In particular, we examine whether there is evidence of an heterogeneous transmission mechanism of the shocks by conditioning on economic and financial conditions of group of regions. We also partially relax the assumption of constant coefficients over time, by investigating the effects of fiscal policy in the sample period before and after the Global Financial Crisis (GFC).

4.3.2 Structural form model and identification.

Following Lütkepohl (2005), the structural form model is described by the following equation linking the reduced form innovations $u_{i,t}$ to the structural shocks $\epsilon_{i,t}$:

$$u_{i,t} = \Gamma \epsilon_{i,t} \tag{4.2}$$

where $\epsilon_{i,t} \sim N(0, I)$, and I is the identity matrix, implying that the structural shocks are orthogonal to each other. Since our aim is to focus on the identification of the effects of government expenditure and tax shocks, we focus on the estimation of the coefficients on the first and second column of the structural form matrix Γ[51]. Previous studies on the Italian regional fiscal multipliers resorted to Cholesky identification (see, e.g., Deleidi et al. 2021; Destefanis et al. 2022)[52].

[49] The software program retains draws, obtained from the algorithm, that satisfy the stability of the VAR (see Appendix A4). The R code is available at the link provided in the section "Software and codes" at the end of this thesis.

[50] We have also applied the hierarchical model proposed by Jarociński (2010), to allow regional heterogeneity in the parameters. However, the hyperparameter that drives the cross-sectional heterogeneity of the coefficients is practically zero, and the algorithm does not converge, even after 1,000,000 of iterations. This corroborates our choice of assuming homogeneity.

[51] As stated in section 4.1, we focus on economy wide style shocks to fiscal policy. Our aim is to study the effects of fiscal policy shocks on the housing market and credit market conditions. These shocks may have different effects and policy implications from policy measures specifically targeted to the housing market (see Egan and Bergin, 2023).

[52] Proxy SVAR is another methodology used in the literature to identify fiscal shocks. See, amongst all, the study of Angelini et al. (2022) based on country-level data for the US disentangling the role payed government spending and tax revenues shocks in shaping output dynamics. To our knowledge, the only study based on the identification of local fiscal shocks (in particular, to government consumption) through proxy SVAR fitted to NUTS2 data is the one by Matarrese and Frangiamore (2023).

Since we have annual data, we want to avoid recursive short run restrictions, which would imply assuming that regional fiscal policy does not respond to shocks hitting the real economic activity, the housing market and the credit market within one year. In this study we use the sign restrictions approach, pioneered by Uhlig (2005) to identify monetary policy shocks. This method consists of imposing reasonable restrictions on the sign of the response of some endogenous variables to the shock of interest, drawing from the economic theory or the empirical evidence[53].

In particular, we follow Canova and Pappa (2007) and Pappa (2009) in setting the restrictions. We assume that, on impact, government spending shocks push government expenditure up, and increases output. The latter assumption is supported by a review of the empirical studies on the effects of government expenditure and on the estimation of government spending multipliers (see Table C4.1 in Appendix C4) that shows that, by employing either country-level and regional-level data for Italy, there is evidence of an increase in output after a positive government spending shock. Furthermore, Pappa (2009), using both RBC and New Keynesian models with sticky prices, shows that a shock to different components of government expenditure leads to an output increase. Finally, the sign restriction used to identify a tax shock are such that, on impact, there is an increase in government revenues and a decrease in the output due to an increasing tax burden on the local economy (Canova and Pappa 2007).

This identification approach has some advantages. As the reader may notice, we do not impose zero restrictions, thus the timing of the response is fully unrestricted, which means that we do not rely on delay restrictions (Pappa 2009), and this solve issues like predictability of fiscal shocks and non-fundamentalness (see, e.g., Ramey 2011; Forni and Gambetti 2010).

Table 2 summarizes our identifying restrictions. The empty spaces indicate that the response of the variables to the shock is left unrestricted.

Table 4.2: Sign restrictions.

	variables					
shocks	G	T	Y	HP	L	i
government spending	+		+			
tax		+	-			

Note: G ≡ gov. exp.; T ≡ gov. rev.; Y ≡ output; HP ≡ house prices; L ≡ loans; i ≡ interest rates.

Therefore, our identification strategy is agnostic with respect to the response of house prices, loans and interest rates. This is because we are interested in the effects of fiscal policy on these variables (Uhlig 2005).

The simulation approach suggested by Dieppe et al. (2016), to implement the sign restriction approach put forward by Uhlig (2005), is described in Appendix B4.

[53]The study of Lucidi (2022) relies on sign restrictions to compute fiscal multipliers for the Italian economy using NUTS-2 data. However his target variable is real output (contrary to our study, where the main focus is on house prices and proxy of credit market conditions, such as loans and interest rates).

4.4 Results

4.4.1 Baseline results

Figure 4.1 shows the IRFs to an expansionary fiscal policy shock obtained from the baseline model, with regional and time dummies, and a linear and quadratic time trend included among the exogenous variables. Panel (a), (b) and (c) show the results obtained from the model using mortgages, loans for durable goods and loans for consumer credit, respectively. The top panel shows the effects of a positive government spending shock, whereas the bottom panel shows the effects of a negative tax shock (the arrows indicate the direction of the causation; for example $G- > Y$ is the IRFs of output to a government spending shock). The IRFs are normalized so that the response of government expenditure to government spending shocks and the response of government revenues to tax shocks is equal to one. This is done to have both spending and tax shocks of equal magnitude to ease the interpretation of the results across different model specifications.

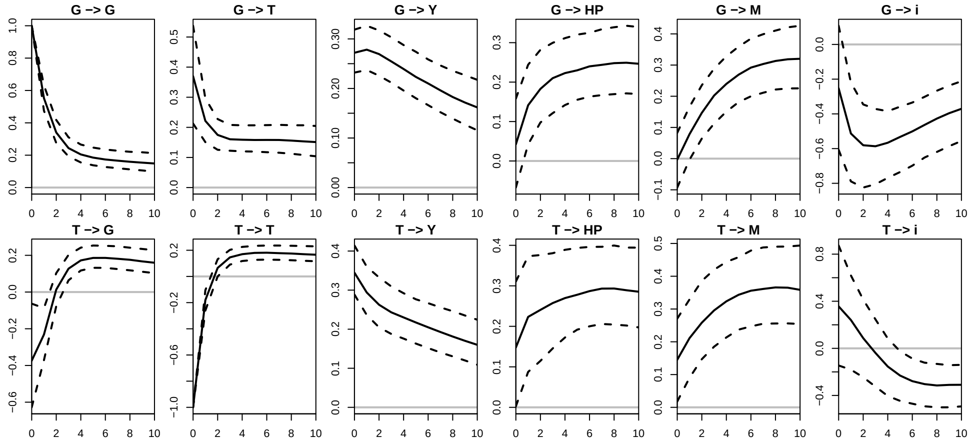
We can notice that the results are very similar across the three panels. Figure 4.1 shows that, on impact, a tax cut stimulates the real economic activity more than government spending increase does, but the magnitude of the output responses to these two shocks converges in the long-run. Indeed, output goes up by about 0.28% and 0.35% on impact, after an increase in spending and a reduction in taxes, respectively. However, while the response to tax shocks start reducing, the one to spending shocks peaks one year after at about 0.29% and then start reducing as well. The positive response is persistent in both cases, and after ten years the effects of both shocks on output converge to each other, being about 0.15%.

As for the volume of loans, we can observe that a government spending increase is found to increase mortgages and, to less extent, the volume of loans for durables and consumer credit. More specifically, mortgages show an increasing response after a government spending impulse, with a long-run increase equal to 0.3%; loans for durables increase significantly after a government spending shock only on impact, by 0.2%; the response of consumer credit is hump-shaped and peaks at 0.15% after six years[54]. A tax cut raises all three categories of loans (more than government spending), especially mortgages, which go up by about 0.35% in the long-run, whereas the response of loans for durables and consumer credit is hump-shaped, reaching a peak of 0.4% and 0.3%, respectively, three years after the tax cut.

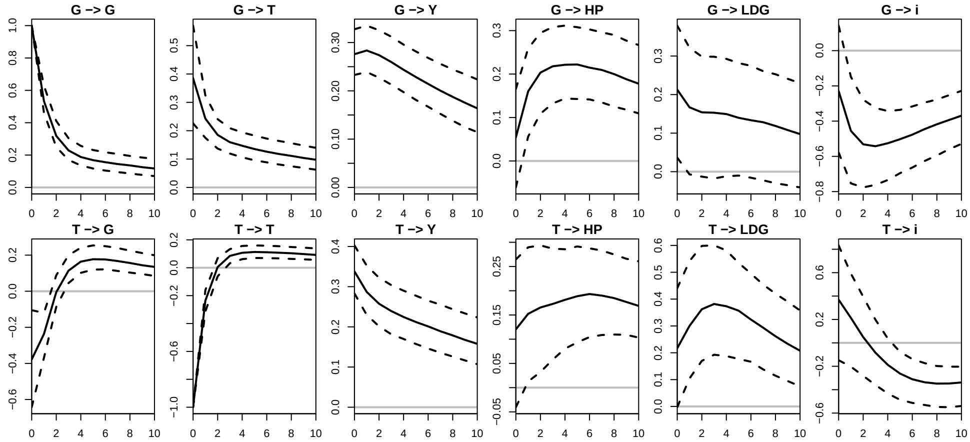
While the response of loans is stronger in case of a tax cut, government spending increases produce a larger fall in the interest rates. In particular, inspection of panel (a), (b) and (c) shows a fall in the loan interest rate in response to a positive government spending shock: the response, on impact, is a 20 basis points decrease, and the peak fall, occurring three years after the shock, is about 60 basis points. In the long-run, the interest rate response is a fall by 40 basis points. These empirical findings are in line with those by Auerbach et al. (2020), who find a decrease in households interest rate after an increase in government (military) expenditure at the local level for the U.S. counties and those by Murphy and Walsh (2022), providing empirical evidence supporting the positive effects of fiscal policy on the credit market conditions in the US.

[54] The increase in household loan volume due to the rise in government spending is also observed by the study of Cipollini and Frangiamore (2023) using Italian NUTS-3 data.

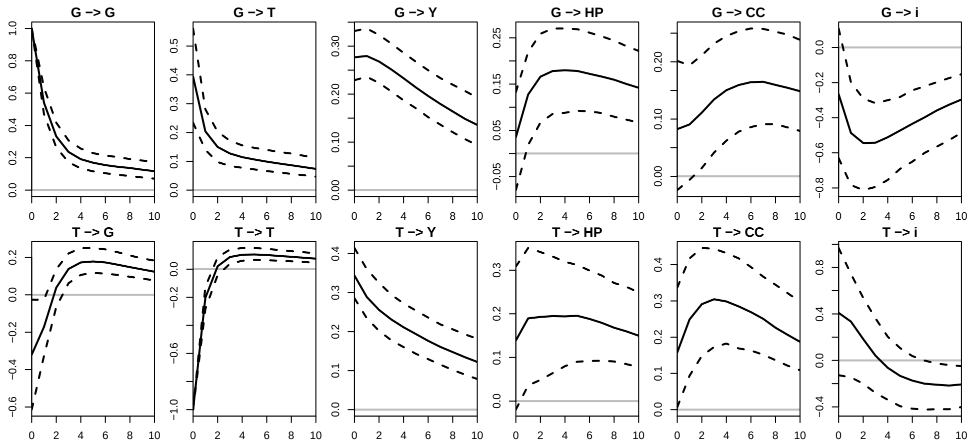
Figure 4.1: Impulse response functions and 68% credibility intervals (baseline models).



(a) model with mortgages



(b) model with loans for durables



(c) model with consumer credit

Note: G ≡ gov. exp.; T ≡ gov. rev.; Y ≡ output; HP ≡ house prices; M ≡ mortgages; LDG ≡ loans for durables; CC ≡ consumer credit; i ≡ interest rates. The solid lines are the posterior medians, the dashed lines are the 68% credibility intervals. The direction of the arrows indicates the direction of the causation. For example, $G- > Y$ is the IRFs of output to a government spending shock. The first line of each panel shows the effects of a one standard deviation increase in government spending, whereas the second panel the effects of a tax cut of the same magnitude. The IRFs measures the elasticity of the endogenous variables to one standard deviation shock in the fiscal variables, being the endogenous variables transformed into log-levels. The horizontal gray line is the line at zero.

The response of the loan interest rate to a tax cut is smaller than the one to a government spending shock: it is (statistically) significant and negative only in the long run.

Finally, an expansionary fiscal policy leads to an increase in house prices. In particular, after a government spending shock, house prices go up by 0.05% on impact, and about 0.15% after one year, and this response is similar across the three panels (a), (b) and (c). As far the response of tax cuts, house prices increase by 0.15% on impact in all the three models. However, the long-run response is different across the three panels: in panel (a) house prices reach a peak response of 0.3% seven years after the shock; the results in panel (b) and (c), obtained from the models with loans for durable goods and consumer credit, show that the peak response of house prices occurs six years and five years after the shock, respectively, and it is equal to 0.2%.

Our findings confirm the empirical evidence for Italy provided by Afonso and Sousa (2011) using country-level data. However, while their results are less clear in terms of the transmission mechanism (given a raise in the average cost of debt and a decrease in output over the long-run), our analysis (based on NUTS-2 data) shows that both the income and the credit channel contribute to the increase in house prices. Moreover, looking at the interaction between the response of house prices and household lending after a fiscal policy expansion, the positive effects on house prices and mortgages can also be explained by taking into account the role of houses used as collateral to obtain new credit, as suggested by Khan and Reza (2017).

4.5 Heterogeneity: do regional characteristics and the period under investigation matter?

In this section we study the effects of fiscal policy on housing prices and credit market conditions for different sub-samples according to a proxy of economic development, or to a proxy of financial development, and, finally, by considering the sub-sample before and after the Global Financial Crisis (GFC), which was a relevant episode for the housing market.

4.5.1 Heterogeneity: the role of economic development

Recently, Gabriel et al. (2023) showed, using NUTS-2 data for the Euro area, that government spending multipliers in the core regions are higher than those belonging to the periphery countries. The striking different level of economic development of Italian regions, with the Mezzogiorno lagging behind, suggests to assess whether a proxy of economic development, such as income per capita, would be an important driver of the heterogeneous impact of fiscal policy shocks on output, credit market conditions and house prices. Therefore, taking the average GDP per capita level of the Italian regions, we split the sample in two parts: (i) one containing the regions with a GDP per capita lower than the median; (ii) the other one including the regions with a GDP per capita higher than the median. The model estimation is based on the within transformation to reduce the number of parameters. Not surprisingly, the bulk of the regions in the first sub-sample are the ones located in the Mezzogiorno[55].

The empirical results are shown in Figure 4.2: the red lines plot the average impulse response to fiscal

[55]We find that the regions with a lower level of GDP per capita are Abruzzo, Basilicata, Calabria, Campania, Marche, Molise, Puglia, Sardegna, Sicilia and Umbria, all belonging to the Mezzogiorno, except for Marche and Umbria, which are considered in the Centre according to the NUTS-1 classification. The other regions, namely, Emilia-Romagna, Friuli-Venezia Giulia, Lazio, Liguria, Lombardia, Piemonte, the autonomous provinces of Bolzano and Trento, Toscana, Valle d'Aosta and Veneto, have a higher GDP per capita.

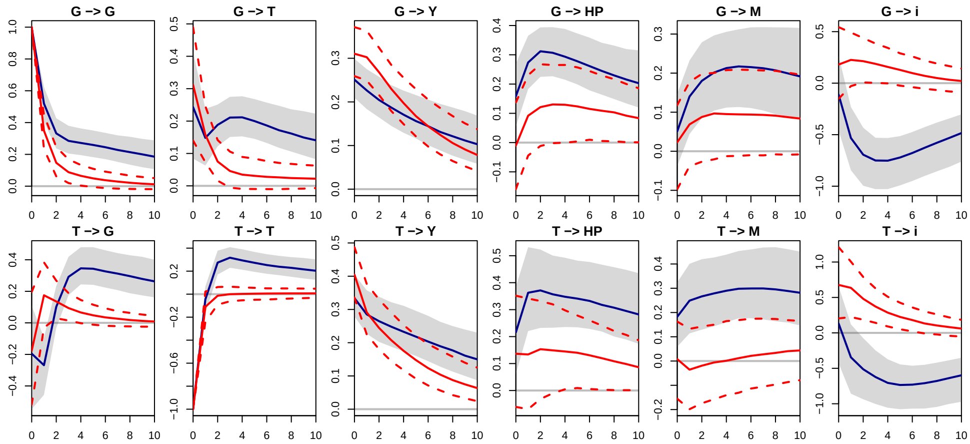
shocks for the group of regions with an income per capita below the median; the blue lines with shaded area plot the average response for the group with an income per capita above the median. The IRFs are normalized in order to have the response of government expenditure and revenues to spending shocks and tax shocks equal to one, respectively.

A first set of results (see panel (a), (b) and (c)) confirms the empirical findings of previous studies related to Italian regions, pointing at lower effects of fiscal policy in the Mezzogiorno macro-region (see, e.g., Deleidi et al. 2021; Lucidi 2022; Matarrese and Frangiamore 2023). In particular, the GDP reacts more to tax shocks in the regions with an income per capita above the median, especially in the long-run, whereas the GDP response to a spending shocks is similar in the two groups of regions.

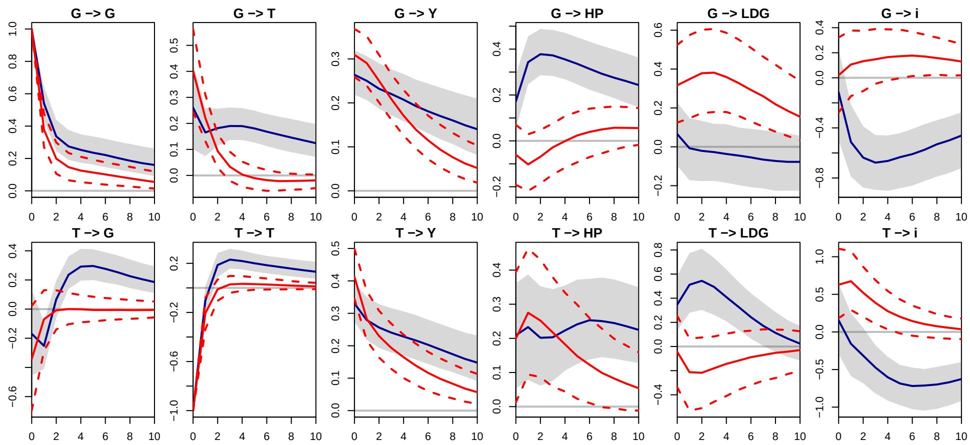
Panel (a) shows the empirical evidence related to the model with mortgages, suggesting that the response of house prices to either government spending or to a tax shock are larger in the more developed regions. In particular, in the regions with income per capita above the median, the house price increase reaches a peak of 0.3% two years after a positive government spending shock, which is three times as much as in the regions with income per capita below the median, where the peak response is 0.1%. The response of house prices to a tax cut is stronger than the one associated with government spending, reaching a peak of 0.35% after two years in the group of regions with income above the median, versus 0.1% peak in the group of regions with income below the median. Moreover, the empirical evidence points at a larger improvement in access to credit market for the more developed regions. In particular, there is evidence of a long-run response in the volume of loans to a government expenditure innovation equal to a 0.2% (0.1%) in the group of regions with income per-capita above (below) the median. A tax cut leads to a stronger increase in mortgages equal to 0.3% in the more developed regions, versus a response close to zero in the less developed ones. The most striking differences are in the response of loan interest rates to both a positive government spending and a tax cut shock, being negative only for the group of regions with income per capita above the median (the interest rate fall is between 50 to 80 basis points). For the other two categories of loans, namely loans for durables and consumer credit, panels (b) and (c) show that a tax cut is more effective in the more developed area, with a peak response of the loans for durable goods equal to about 0.6%, against a negative peak (which, however, is not statistically significant) equal to about -0.2% in the less developed regions, and a peak response of consumer credit equal to 0.3% in the regions with higher GDP per capita, against a 0.2% in the regions with a lower GDP per capita. The responses of the other endogenous variables in the models involving these two categories of loans are similar to those shown in panel (a), obtained from the model that includes mortgages, where, essentially, house prices rise and interest rates fall more in regions with higher per capita income.

In summary, the effectiveness of an expansionary fiscal policy on house prices and credit market conditions prevail in the more developed regions, especially through a tax cut easing access to credit.

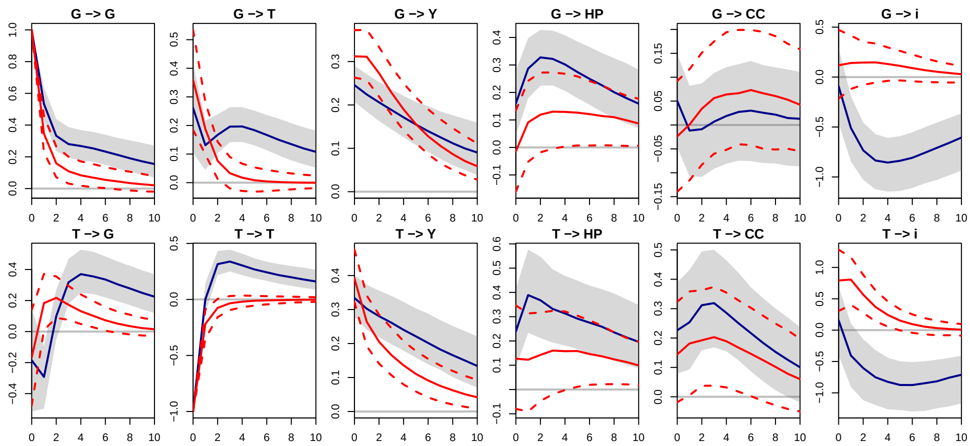
Figure 4.2: Impulse response functions and 68% credibility intervals in less developed (red lines) and more developed (blue lines with shaded area) regions.



(a) model with mortgages



(b) model with loans for durables



(c) model with consumer credit

Note: G ≡ gov. exp.; T ≡ gov. rev.; Y ≡ output; HP ≡ house prices; M ≡ mortgages; LDG ≡ loans for durables; CC ≡ consumer credit; i ≡ interest rates. The solid lines are the posterior medians, the dashed lines and the shaded areas are the 68% credibility intervals. The direction of the arrows indicates the direction of the causation. For example, $G->Y$ is the IRFs of output to a government spending shock. The first line of each panel shows the effects of a one standard deviation increase in government spending, whereas the second panel the effects of a tax cut of the same magnitude. The IRFs measures the elasticity of the endogenous variables to one standard deviation shock in the fiscal variables, being the endogenous variables transformed into log-levels. The horizontal gray line is the line at zero.

4.5.2 Heterogeneity: the role of financial development of the regions

In this section we investigate whether, conditioning on the level of financial development of different group of regions, plays a role in the transmission mechanism of fiscal policy shock to house prices and credit market conditions. In line with Rossi and Scalise (2022) we use, as a proxy of financial development at local level, the number of bank branches per inhabitant[56]. In particular we take the information on the number of bank branches per 100,000 inhabitants from the Bank of Italy and we split the sample in two: (i) regions with a number of bank branches per inhabitant less than the median; (ii) regions with a number of bank branches per inhabitant higher than the median[57].

We use the within transformation of the data for each sub-sample and the red (blue) line plots in Figure 4.3 shows the impulse responses to fiscal shocks associated with the regions with a low (high) level of banking sector development, whereas the dashed lines (the gray shaded area) are the 68% credibility intervals.

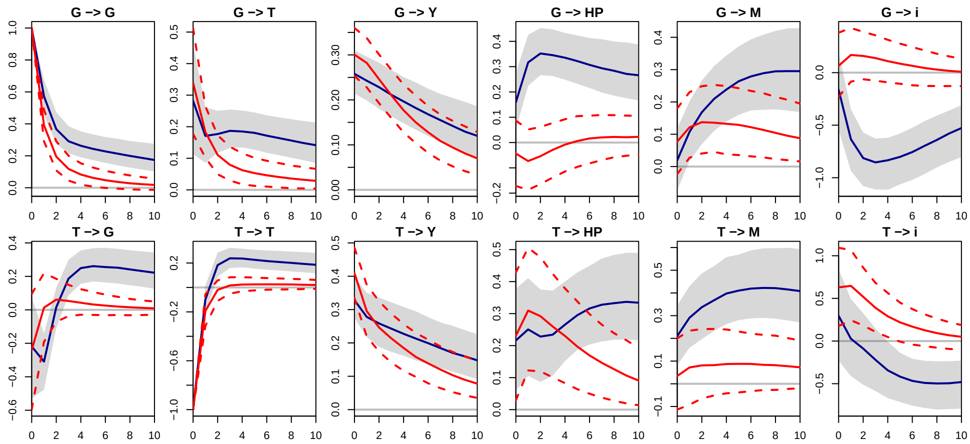
The plots in Figure 4.3 show that fiscal policy is more effective in regions with a higher level of banking sector development. In particular, the effects of positive government expenditure and tax cut shocks on output are larger (over a long-run horizon) in the regions with a higher level of banking sector development, regardless of the category of loans considered in the model specification. An expansionary fiscal policy benefits the access to credit market (especially through an interest rate fall) in the regions with a higher level of banking sector development. In particular, there is evidence of a negative reaction (with a negative peak equal to 80 basis points) to a positive government spending shock only in the regions with a higher level of banking sector development. Only the regions more financially developed benefit from an interest fall due a tax cut (while those less financially developed experience the opposite). In particular, the reduction in interest rates, as a response to a tax cut, reaches 50 basis points in the long-run when considering the model with mortgages, it peaks at 80 basis points and converges to 50 basis points in the long-run in the case of the model with loans for durables, and reaches 60 basis points when considering the model with consumer credit.

The response of loan volume to a government spending shock is different according to the category considered: while the positive effect on mortgages is stronger in the more financially developed regions, there is evidence of a rise in the volume of loans for durables and consumer credit in the regions with a lower level of banking sector development. Finally a tax cut is more beneficial to the most financially developed regions given the evidence of a rise in the volume of loans (mortgages and those used for buying durables).

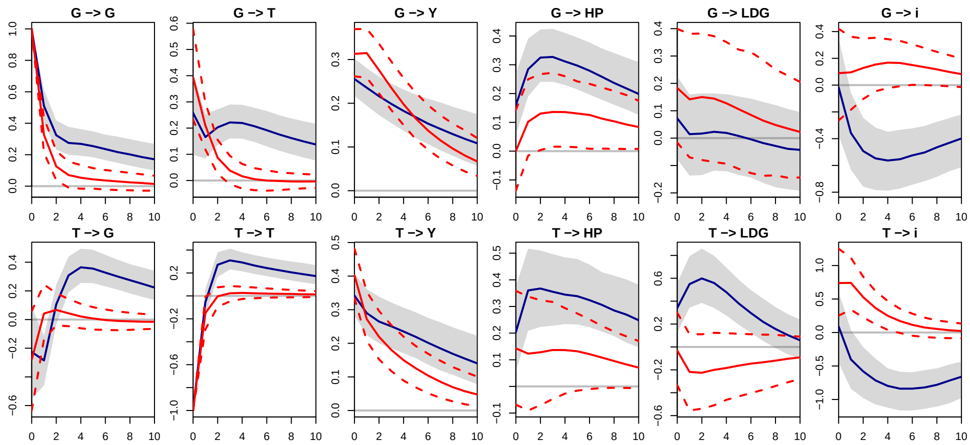
[56]See also the study of Destefanis et al. (2022) on fiscal multipliers and the role played by banking sector development

[57]Since data availability is from 2008, we compute, for each Italian region, the average of the number of bank branches per 100,000 inhabitants over the period 2008-2019. The Italian regions with a number of bank branches per inhabitants lower than the median are: Abruzzo, Basilicata, Calabria, Campania, Lazio, Liguria, Molise, Puglia, Sardegna and Sicilia; the ones with a higher number of bank branches per inhabitants are: Emilia-Romagna, Friuli-Venezia Giulia, Lombardia, Marche, Piemonte, Toscana, the two autonomous provinces of the Trentino-Alto Adige region, Umbria, Valle d'Aosta and Veneto.

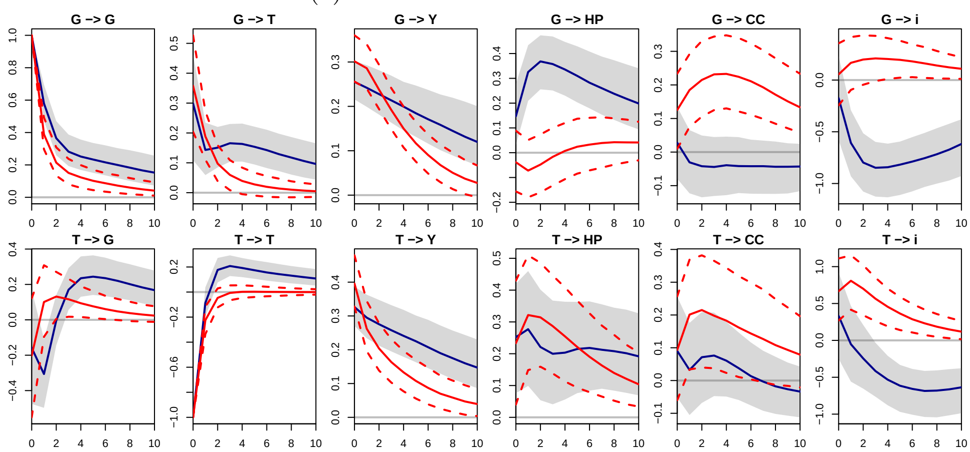
Figure 4.3: Impulse response functions and 68% credibility intervals in regions with low level of financial development (red lines) and in regions with high level of financial development (blue lines with shaded area) regions.



(a) model with mortgages



(b) model with loans for durables



(c) model with consumer credit

Note: G ≡ gov. exp.; T ≡ gov. rev.; Y ≡ output; HP ≡ house prices; M ≡ mortgages; LDG ≡ loans for durables; CC ≡ consumer credit; i ≡ interest rates. The solid lines are the posterior medians, the dashed lines and the shaded areas are the 68% credibility intervals. The direction of the arrows indicates the direction of the causation. For example, $G- > Y$ is the IRFs of output to a government spending shock. The first line of each panel shows the effects of a one standard deviation increase in government spending, whereas the second panel the effects of a tax cut of the same magnitude. The IRFs measures the elasticity of the endogenous variables to one standard deviation shock in the fiscal variables, being the endogenous variables transformed into log-levels. The horizontal gray line is the line at zero.

There is a striking difference in the response of house prices to government spending shocks: there is no impact in regions with a low level of banking sector development, while, for the more financially developed regions, the impact on house prices is statistically significant and positive (reaching a peak of 0.35% two years after the shock and a long-run response of about 0.3%). A tax cut has a larger positive effect on house prices in the group of regions with a higher level of financial development (especially when considering a model with mortgages and loans for durables, which is three times larger).

To summarize, there is evidence that fiscal policy is more effective in regions with a higher level of banking sector development, especially when we consider the interplay between the market of mortgages and house prices.

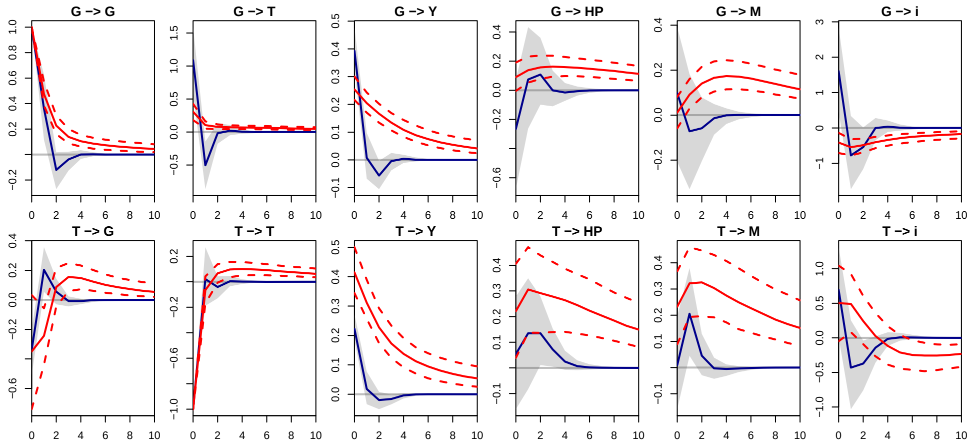
4.5.3 Before and after the Global Financial Crisis

A number of recent studies investigate whether fiscal policy can be a valid additional tool to help the recovery from the Global Financial Crisis, capturing a relevant time span of the sample period examined in this study. The study of Amendola et al. (2020), focusing on the Euro Area, finds that fiscal policy is more effective during the Zero Lower Bound period characterizing short-term interest rates. In this study we also assess whether the reaction of GDP, house prices and credit market conditions to fiscal policy shocks has changed after the GFC. For this purpose, we estimate the Panel VAR using two sub-samples: (i) the first one involves the period 2004-2008, and we label this the period before the GFC[58]; (ii) the second one is from 2009 up to 2019. Figure 4.4 shows the results for the three models: panel (a) refers to the model with mortgages; panel (b) the one with the loans for durable goods and panel (c) the one with consumer credit. The solid blue lines are the responses in the period before the GFC, with the shaded area being the associated 68% credibility intervals, whereas the red solid lines are the responses in the period after the GFC, with the corresponding 68% credibility intervals represented by the red dashed lines.

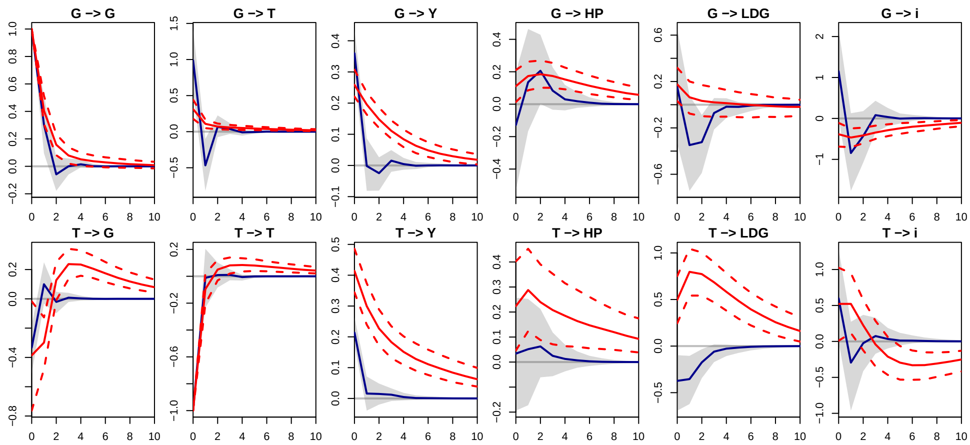
Figure 4.4 shows that, in line with Amendola et al. (2020), fiscal policy effects on GDP are larger after the GFC: there is a positive response to either an increase in government spending or to a tax cut with an impact twice as large as the one associated with innovations to expenditure. Moreover, fiscal policy acts as a stimulus to house prices, especially through a tax cut, only after the GFC. The increase in house prices following a tax cut shock is 0.2% after the GFC, more than twice as large as that recorded before the GFC and with a peak response of about three times that recorded in the pre-GFC period. The response of house prices to a government spending increase is larger after the GFC, however with a similar peak response, but with a larger and more persistent response in the long-run. Finally, a fiscal policy shock relaxes credit market conditions only after the GFC: there is evidence of an interest rate fall only in response to a government spending increase, while the reaction to a tax cut is statistically insignificant. More specifically, interest rates fall by about 50 basis points after the GFC in reaction to a positive government spending shock, whereas the response is not significant in the pre-GFC period. As for tax cut shocks, there is a long-run improvement in the cost of credit, in the model with mortgages and loans for durables only, where the interest rates go down by about 25 basis points.

[58] Our data are annual, thus, since the Lehman bankruptcy started in September 2008, we included the 2008 in the first sub-sample.

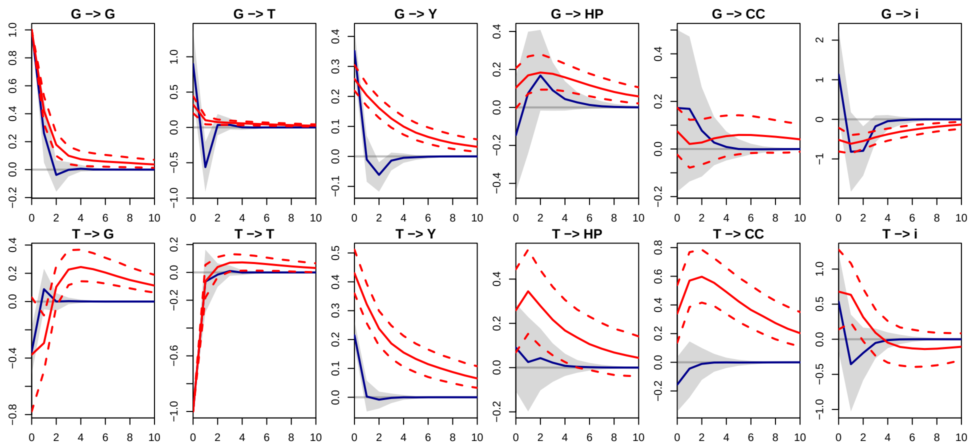
Figure 4.4: Impulse response functions and 68% credibility intervals before (blue lines with shaded area) and after (red lines) the GFC.



(a) model with mortgages



(b) model with loans for durables



(c) model with consumer credit

Note: G ≡ gov. exp.; T ≡ gov. rev.; Y ≡ output; HP ≡ house prices; M ≡ mortgages; LDG ≡ loans for durables; CC ≡ consumer credit; i ≡ interest rates. The solid lines are the posterior medians, the dashed lines and the shaded areas are the 68% credibility intervals. The direction of the arrows indicates the direction of the causation. For example, $G- > Y$ is the IRFs of output to a government spending shock. The first line of each panel shows the effects of a one standard deviation increase in government spending, whereas the second panel the effects of a tax cut of the same magnitude. The IRFs measures the elasticity of the endogenous variables to one standard deviation shock in the fiscal variables, being the endogenous variables transformed into log-levels. The horizontal gray line is the line at zero.

Moreover, while either a tax cut or an increase in fiscal spending raises mortgages, the volume of loans to buy durables and consumer credit are positively affected only by a tax cut. In particular, while, before the GFC, the response of mortgages to positive government spending and tax cut shocks is either insignificant or very small, after the GFC, the peak level in the response of mortgages to a positive government spending shock is equal to 0.2% and the one associated with a tax cut is equal to 0.3%. The response of loans for durable goods and consumer credit is statistically significant and positive after the GFC only after a tax cut, whereas the responses of these two categories of loans to government spending shocks is not significant in neither of the two regimes. In particular, after the GFC, a tax cut raises loans for durables by 0.5% on impact, and then this response reaches a peak at about 0.8% after one year (whereas the short-run response is lower and even negative before the GFC). A similar reaction is observed in the period after the GFC for the consumer credit: the response to a tax cut is equal to 0.4% on impact and it peaks at 0.6% two years after the shock.

Overall, we find that fiscal policy, during and after the Global Financial Crisis, has played a counter-cyclical role, stimulating the housing market and relaxing the credit market conditions (in line with the studies of Auerbach et al. 2020 and Murphy and Walsh 2022), hence complementing monetary policy.

4.6 Conclusions

In this study we focus on the effects of fiscal policy on house prices and credit market conditions (loans and interest rates) for the households in Italy. The literature, so far, has mainly concentrated on country-based studies and, to our knowledge, this is the first study using regional data at the NUTS-2 level. Moreover, we disentangle the contribution of shocks to government spending and to tax revenues in shaping the dynamics of house prices and credit market conditions. We use a Panel VAR fitted to 21 regions (or 20 regions in the case of the consumer credit) over the period 2004-2019. The reduced-form model is estimated by employing Bayesian estimation procedures, and the identification of the two fiscal shocks is obtained through sign restrictions. The empirical evidence suggest that, although an expansionary fiscal policy can provide a stimulus to an house appreciation and it can ease the access to credit, the positive effect is only on the group of regions with an income per capita or an index of banking sector development (bank branches per inhabitant in a given region) above the median. Finally, the empirical findings show that fiscal policy played a counter-cyclical role over the Global Financial Crisis and Sovereign Debt Crisis time period.

These results have important policy implications, as they inform on the potential counter-cyclical role of fiscal policy in stimulating the housing and credit markets, thus providing support for the complementarity between fiscal and monetary policy. However, the analysis has some limitations that need to be acknowledged. First, since we focus on broad fiscal shocks, and thus do not delve into different types of fiscal changes, particularly those targeted at the housing market or affecting housing taxation, we are unable to capture the relevant institutional changes in the Italian tax system that have occurred over the last two decades. Our empirical approach and the regional public revenue information we use are not suitable to study these important issues and to identify regional heterogeneity in the response to these specific shocks. Second, we only derive an average elasticity, which indicates the effects of fiscal policy changes in one region relative to another, on the relative conditions of housing and credit markets. We do not consider spillover effects between regions, which might be relevant in the Italian regional context. All these observations can be a stimulus for future research.

Appendices to chapter 4

Appendix A4: Bayesian estimation

The model in (4.1) can be rewritten in a more suitable way for the discussion of the Bayesian estimation, by stacking all the time observations of every unit, and combining all the regressors, $Y_{i,t-1}$ and $W_{i,t}$, in a matrix X. The model is now as follows:

$$Y = XB + U \tag{4.3}$$

where Y is the $NT \times n$ matrix of endogenous variables, X is the $NT \times (np+m)$ matrix of regressors, with an associated $(np+m) \times n$ matrix of coefficients B, and U is the $NT \times n$ matrix of reduced-form residuals, in stacked form[59]. As stated in section 4.3.1, we follow the approach of Banbura et al. (2007), using dummy observations (or artificial data), to impose a Normal-Wishart prior. This method consists of incorporating prior information into the VAR, by mixing the actual data with artificial data constructed in a proper way, to satisfy the prior assumptions one wishes (Blake and Mumtaz, 2012). Let us consider the representation in (4.3), and suppose that we construct Y_D and X_D artificial observations. The Normal-Inverse Wishart prior is represented by the following:

$$\begin{aligned} &p(B|\Sigma) \sim N\left(b_0, \Sigma \otimes (X_D' X_D)^{-1}\right) \\ &p(\Sigma) \sim IW\left(S, T_D - n + m\right) \end{aligned} \tag{4.4}$$

with prior moments given by:

$$\begin{aligned} &B_0 = (X_D' X_D)^{-1} X_D' Y_D \\ &b_0 = vec(B_0) \\ &S = (Y_D - X_D B_0)'(Y_D - X_D B_0) \end{aligned} \tag{4.5}$$

and T_D equal to the number of artificial observations. The conditional posterior distributions are as follows:

$$\begin{aligned} &H(b|\Sigma, Y) \sim N\left(vec(B^*), \Sigma \otimes (X'^* X^*)^{-1}\right) \\ &H(\Sigma|b, Y) \sim IW\left(S^*, T^*\right) \end{aligned} \tag{4.6}$$

with T^* the degrees of freedom, equal to the total number of observations with the artificial data appended, and S^* the scale matrix of the Inverse-Wishart distribution:

$$\begin{aligned} &B^* = (X'^* X^*)^{-1} X'^* Y^* \\ &S^* = (Y^* - X^* b)'(Y^* - X^* b) \end{aligned} \tag{4.7}$$

[59]N is the number of regions (21 or 20); T is the time series length (15), n is the number of endogenous variables (6); p is the number of lags (1); m is the number of exogenous variables (depending on the specification).

where $Y^* = \begin{bmatrix} Y \\ Y_D \end{bmatrix}$ and $X^* = \begin{bmatrix} X \\ X_D \end{bmatrix}$ (see Blake and Mumtaz, 2012).

The dummy observations to implement the prior in (4.5) are constructed as in Banbura et al. (2007), in order to match the Minnesota moments (Litterman, 1986):

$$Y_D = \begin{pmatrix} \mathrm{diag}(\delta_1\sigma_1, \cdots, \delta_n\sigma_n)/\lambda \\ 0_{n(p-1)\times n} \\ \cdots \\ \mathrm{diag}(\sigma_1, \cdots, \sigma_n) \\ \cdots \\ 0_{m\times n} \end{pmatrix} \qquad X_D = \begin{pmatrix} J_p \otimes \mathrm{diag}(\sigma_1, \cdots, \sigma_n)/\lambda & 0_{np\times m} \\ \cdots & \cdots \\ 0_{n\times np} & 0_{n\times m} \\ \cdots & \cdots \\ 0_{m\times np} & \mathrm{diag}(\epsilon)_{m\times m} \end{pmatrix} \tag{4.8}$$

where δ_i are the prior means of the first lag of the i-th endogenous variable, set equal to 1 in the spirit of the Minnesota prior, to incorporate the information that log-level macroeconomic variables are random walk with a unit root; the parameters σ_i are estimated by the standard deviation of the residual of an AR(1) fitted to each endogenous variable, as empirical papers typically do; $J_p = \mathrm{diag}(1, \cdots, p)$, being p the number of lags; the parameter λ controls the overall tightness of the prior, ϵ controls the tightness of the priors on the m exogenous variables, and they are set equal to 1 and 1/1000, respectively, to reflect a loose prior. The first block of the matrices in (4.8) imposes prior information on the coefficients of lags, the second one on the reduced-form covariance matrix, and the last one on the coefficients of the exogenous variables. To simulate the conditional posterior distributions, we implement a Gibbs sampler, performing 5000 repetitions, discarding the first 4000 draws, thus using the last 1000 draws for inference. The algorithm consists of the following steps:

- draw the vectorized matrix of VAR parameters from $H(b|\Sigma, Y)$ in (4.6). In this step, only stable draws are retained, taking the ones whose eigenvalues of the companion matrix are less than or equal to one;
- use the draw of b from the previous step to compute S^*, then draw the covariance matrix of the reduced-form residuals from $H(\Sigma|b, Y)$ in (4.6).

For each iteration, we save the draws of the VAR coefficients and the covariance matrix. Since the Gibbs sampling produces auto-correlated draws, we assess the convergence computing the inefficiency factor for every VAR parameters and the unique elements of the reduced-form covariance matrix, on the last 1000 draws that we retain. The inefficiency factor is given by the following: $IF = 1+2\sum_{i=1}^{20} \hat{\rho}_i$, where $\hat{\rho}_i$ is the i-th order auto-correlation. If the value of the inefficiency factors is around or below 20, it is considered satisfactory (see Primiceri 2005). The following figures show the inefficiency factors for the baseline model (the one with the regional and time dummies) associated with the models that vary according to the category of household loans considered. They are very low, with values around 3 for the VAR coefficients and 3.5 for the unique elements of the variance-covariance matrix.

Figure A4.1: Inefficiency factors of the Panel VAR model with mortgages.

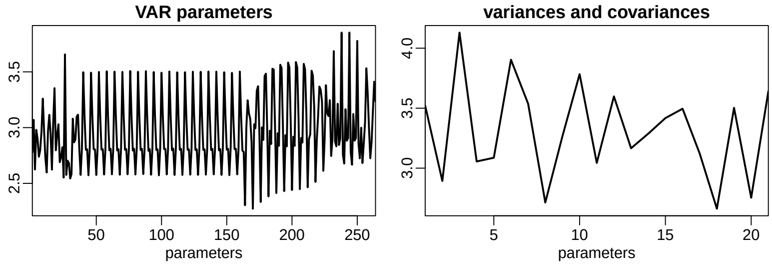


Figure A4.2: Inefficiency factors of the Panel VAR model with loans for durable goods.

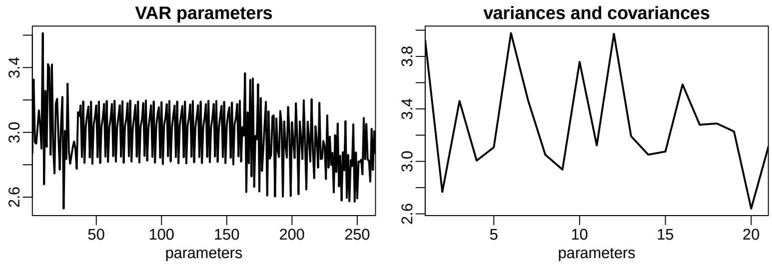
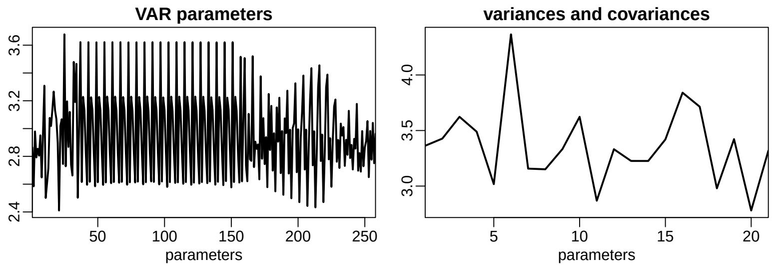


Figure A4.3: Inefficiency factors of the Panel VAR model with consumer credit.



Appendix B4: impulse response functions and sign restrictions

First, we obtain IRFs to reduced-form innovations following the simulation suggested in Dieppe et al. (2016). We start considering the Panel VAR in equation (4.1) at its long-run value (hence the shocks are set equal to zero). Then, a shock hitting a specific variable occurs at time t, and the effects of the shock will propagate to subsequent periods. This accounts to set $Y_{j,T} = 1$ for j equal to the position of the variable to be shocked. In our case, $j = 1, 2$ (government expenditure and government revenues). Using draws of the reduced-form parameters obtained from the algorithm in Appendix A4, we can recover the response of the variables to the shocks, by constructing the values through a recursive algorithm: $\tilde{Y}_{T+1} = \Phi Y_T$, then move one-step ahead using the values from the previous step, $\tilde{Y}_{T+2} = \Phi \tilde{Y}_{T+1}$ and go on until the last horizon H is reached, $\tilde{Y}_{T+H} = \Phi \tilde{Y}_{T+H-1}$. This is done for each retained draw from the Gibbs sampling in Appendix A4. We end up with a posterior distribution of IRFs to innovations in government expenditure and taxes.

Once we obtain the IRFs in reduced form, we implement the algorithm to impose sign restrictions in order to recover the structural IRFs.

The algorithm is described in Dieppe et al. (2016), and it is based on the results of Arias et al. (2014), who show that, using this methodology, it is possible to obtain draws from the posterior distribution of the SVAR model using the draws from the posterior distributions of the reduced-form VAR, by adding an additional orthogonalisation step.

Let us call the matrices of reduced-form IRFs, obtained from the recursive algorithm just described, by Ψ_h. The steps are as follows for each k-th iteration of the Gibbs sampling:

1. Compute the (lower triangular matrix) Cholesky factorization P of the reduced-form covariance matrix $\Sigma = PP^{'}$, and post multiply the reduced-from IRFs by this matrix: $\bar{\Psi}_h = \Psi_h P$;

2. take an orthogonal matrix Q. Define $D = PQ$. This matrix preserve the Structural VAR property: $D[var(\epsilon)]D^{'} = DID^{'} = PQQ^{'}P^{'} = PIP^{'} = PP^{'} = \Sigma$. This is done by drawing a $n \times n$ random matrix M from a multivariate standard normal distribution, $N(0, I)$, and then taking the QR decomposition of this matrix, $M = QR$, where Q is an orthogonal matrix and R is an upper triangular matrix. Then, compute the candidate draws of IRFs $\tilde{\Psi}_h = \bar{\Psi}_h Q = \Psi_h PQ$.

3. check, at the horizon to which the restrictions apply (e.g. $h = 0$ in our model specification), whether the candidate draw satisfies the sign restrictions and discard it only if this does not hold, then repeat step 2.

4. step 2 to 3 is repeated 100 times, thus we end up with 100 valid draws that satisfy the sign restrictions and we save the median over these 100 draws (the results remain the same if we take the mean or the IRFs closest to the median over these 100 draws; these results are available upon request). This will represent the structural IRFs at the k-th iteration of the retained Gibbs draws.

At the end, we get the posterior distribution of the IRFs to make inference. We plot the median IRFs and the 16-th and 84-th quantile as 68% credibility intervals.

Appendix C4: additional tables

Table C4.1: Empirical literature on the effects of government expenditure and on government expenditure multipliers in Italy. The table shows the sign of the effects of government expenditure on output, found by these papers. We use these findings as empirical evidence to impose the restriction that the response of output to government expenditure shock is positive on impact.

Papers	Sign
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.	+
Caprioli, F., & Momigliano, S. (2011). The effects of fiscal shocks with debt-stabilizing budgetary policies in Italy. Bank of Italy Temi di Discussione (Working Paper) No, 839.	+
Deleidi, M. (2022). Quantifying multipliers in Italy: does fiscal policy composition matter?. Oxford Economic Papers, 74(2), 359-381.	+
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Matarrese, M. M., & Frangiamore, F. (2023). Italian local fiscal multipliers: Evidence from proxy-SVAR. Economics Letters, 111185.	+

Appendix D4: Robustness checks

We perform some robustness exercises to check the validity of the results obtained from the baseline VAR.

First, we check whether the empirical findings are different by controlling unobserved heterogeneity using the within transformation of the data. More specifically, we remove the time and regional dummies from the set of exogenous variables in equation (4.1), still keeping the linear and quadratic time trend, and transforming the endogenous variables in the following way: $y_{i,t}^{within} = y_{i,t} - \bar{y}_i - \bar{y}_t + \bar{y}$, that is, we remove the cross-sectional and time averages from the data and add the total average. The within transformation allows to reduce the number of parameters, making the algorithm much faster.

In a second robustness check exercise, the time dummies are replaced by nation-wide control variables to control for the effects of common shocks. We include long-term and short-term interest rates and the price level, to control for monetary policy, which is one of the main confounding factors in the analyses of the effects of fiscal policy, and the oil price Brent to control for global shocks. In Figure D4.1 we compare the results obtained from the baseline models (black lines), with the one obtained from the first exercise (within transformation - blue lines) and from the second exercise (national controls - red lines).

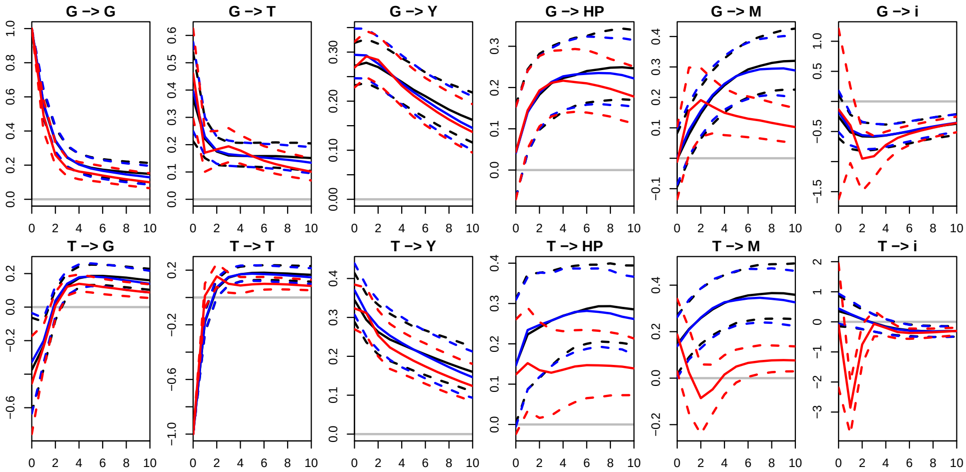
First, we can notice that there is almost no difference between the baseline models and the use of the within transformation[60]. Second, there are major differences between baseline (or within transformation) model and the one that uses nation-wide controls. In particular, the use of country-based control variables (instead of time dummies) leads to a different (although qualitatively similar) response related to the access to credit. Finally, when using the model specification with nationwide control variables, the response of house prices to both type of fiscal shocks is positive but smaller than the one associated with a model with time fixed effects.

Overall, the use of a parsimonious model in terms of the number of parameters to be estimated when using either the within transformation or the use of country based exogenous variables instead of time dummies, confirms that an expansionary fiscal policy can stimulate house prices, thus increasing the value of the collateral, and relax credit market conditions, so reducing the financial constraints and increasing the volume of loans that households obtain for different purposes.

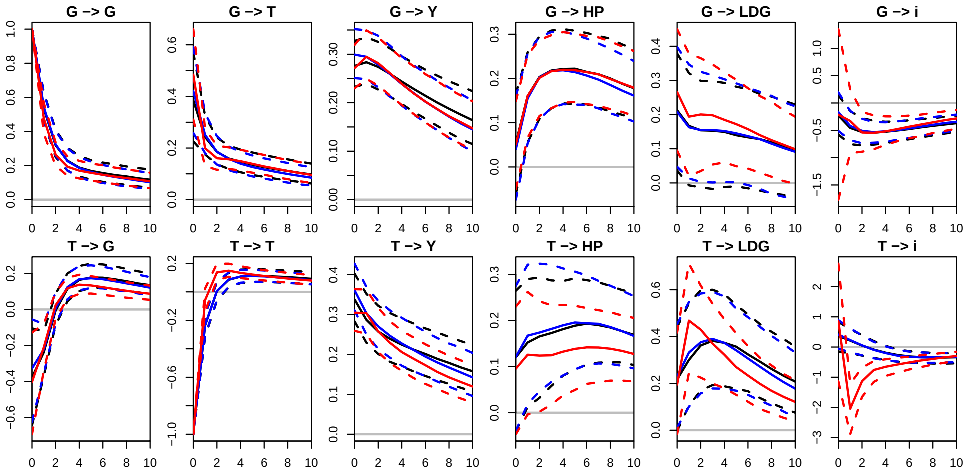
We have also conducted two other robustness check exercises. One is based on the use of either the minimum or the maximum prices of houses provided by OMI database, instead of using the mean between the two. The results are provided in Figure D4.2, where the black lines represents the results of the baseline model with the mean of house prices, whereas the red lines and the blue lines show the results obtained by using the maximum and the minimum, respectively. The other one in which we express government expenditure, government revenues, GDP and loans in per capita terms, and the empirical findings are shown in Figure D4.3. Overall, we can conclude that the results are pretty much similar to the baseline model.

[60]This result has been used in section 4.5, in order to split our sample to check whether the results change according to regional characteristics and the period under consideration, since the within transformation allows to reduce significantly the number of parameters to estimate.

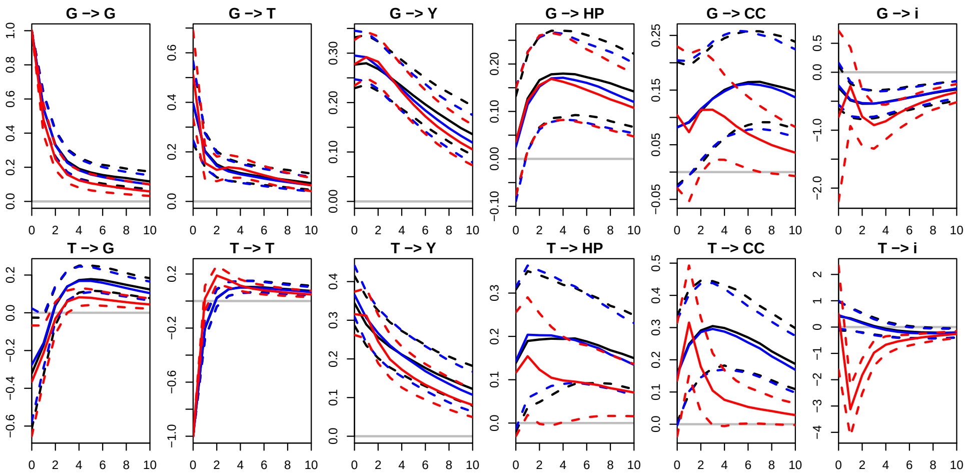
Figure D4.1: Impulse response functions and 68% credibility intervals obtained from three models: (i) in black, the baseline model with regional and time dummies; (ii) in blue, within transformation of the endogenous variables; (iii) in red, replacing the time dummies with country based exogenous variables.



(a) model with mortgages



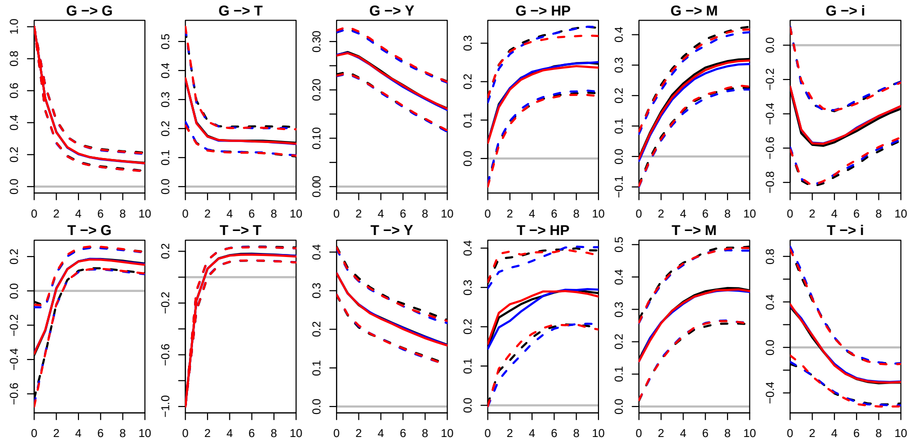
(b) model with loans for durables



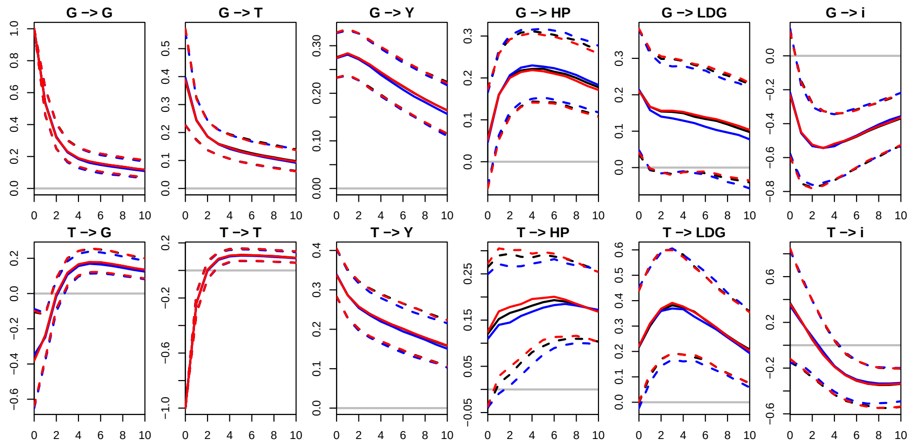
(c) model with consumer credit

Note: G ≡ gov. exp.; T ≡ gov. rev.; Y ≡ output; HP ≡ house prices; M ≡ mortgages; LDG ≡ loans for durables; CC ≡ consumer credit; i ≡ interest rates. The solid lines are the posterior medians, the dashed lines are the 68% credibility intervals. The direction of the arrows indicates the direction of the causation. For example, $G->Y$ is the IRFs of output to a government spending shock. The first line of each panel shows the effects of a one standard deviation increase in government spending, whereas the second panel the effects of a tax cut of the same magnitude. The IRFs measures the elasticity of the endogenous variables to one standard deviation shock in the fiscal variables, being the endogenous variables transformed into log-levels. The horizontal gray line is the line at zero.

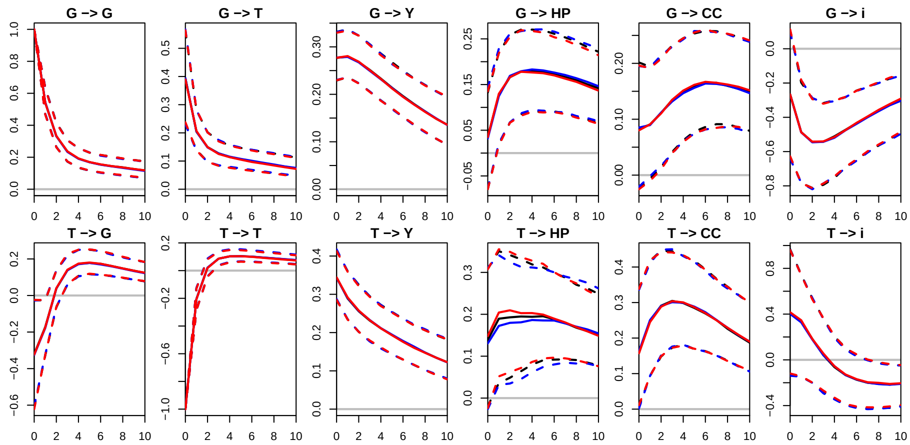
Figure D4.2: Impulse response functions and 68% credibility intervals obtained by using the baseline model (black) with the mean between the minimum and the maximum value of houses, the model with the minimum value of houses (in blue) and the model with the maximum value of houses (in red) provided by OMI.



(a) model with mortgages



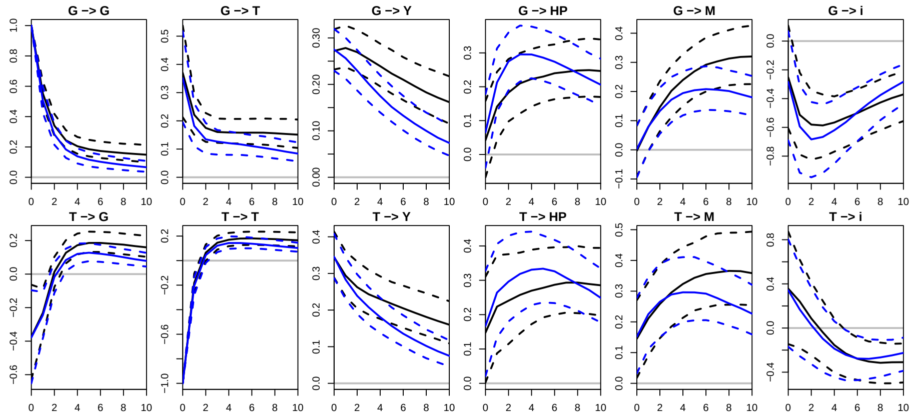
(b) model with loans for durables



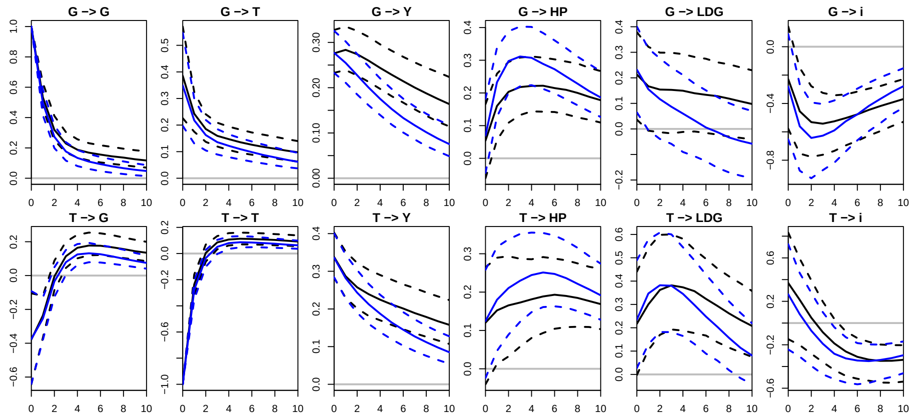
(c) model with consumer credit

Note: G ≡ gov. exp.; T ≡ gov. rev.; Y ≡ output; HP ≡ house prices; M ≡ mortgages; LDG ≡ loans for durables; CC ≡ consumer credit; i ≡ interest rates. The solid lines are the posterior medians, the dashed lines are the 68% credibility intervals. The direction of the arrows indicates the direction of the causation. For example, $G- > Y$ is the IRFs of output to a government spending shock. The first line of each panel shows the effects of a one standard deviation increase in government spending, whereas the second panel the effects of a tax cut of the same magnitude. The IRFs measures the elasticity of the endogenous variables to one standard deviation shock in the fiscal variables, being the endogenous variables transformed into log-levels. The horizontal gray line is the line at zero.

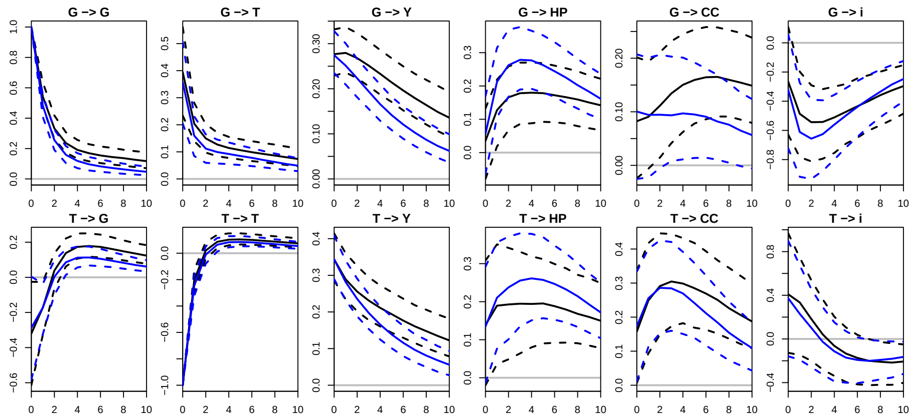
Figure D4.3: Impulse response functions and 68% credibility intervals obtained by using the baseline model (black) and the model with variables in per capita terms (blue).



(a) model with mortgages



(b) model with loans for durables



(c) model with consumer credit

Note: G ≡ gov. exp.; T ≡ gov. rev.; Y ≡ output; HP ≡ house prices; M ≡ mortgages; LDG ≡ loans for durables; CC ≡ consumer credit; i ≡ interest rates. The solid lines are the posterior medians, the dashed lines are the 68% credibility intervals. The direction of the arrows indicates the direction of the causation. For example, $G->Y$ is the IRFs of output to a government spending shock. The first line of each panel shows the effects of a one standard deviation increase in government spending, whereas the second panel the effects of a tax cut of the same magnitude. The IRFs measures the elasticity of the endogenous variables to one standard deviation shock in the fiscal variables, being the endogenous variables transformed into log-levels. The horizontal gray line is the line at zero.

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Concluding remarks

In this thesis I analyse the effects of fiscal policy shocks on output, credit market conditions of different categories of borrowers and housing markets at the regional level in Italy. In particular, I concentrate on different identification schemes implemented using either NUTS-2 or NUTS-3 data to retrieve exogenous fiscal policy shocks. Overall, the results indicate positive effects of fiscal policy on the economy, producing a reduction in credit risk and an improvement in credit market conditions, as well as an expansion of the housing market. However, regional disparities between Northern and Southern Italy shape the heterogeneous impact of fiscal policy, with the positive effects prevailing in the more developed area of the country.

In particular, in the first chapter I show that public spending expansions stimulate the volume of loans in the Italian provinces during the post-Lehman period and the sovereign debt crisis (2011-2018) and also reduce credit risk. This result holds for all borrowers, but also for the non-financial corporate sector, small businesses and households, although the increase in loans to small businesses is smaller than that found for NFCs. These positive effects prevail in the Centre-North and, in particular, the expansion of lending in a macro area after an increase in public spending is found when credit is provided by local banks, especially in the South, thus demonstrating that local fiscal spending does not mitigate the 'home bias' originated by the banking consolidation process that began in the 1990s.

In the second chapter, I study the effects of government spending financed by the European Regional Development Fund (ERDF) on various proxies of regional output. Using a panel dataset at the NUTS-2 level, over the period 1988-2018, I find positive multipliers of ERDF expenditure for GDP, GVA and private sector GVA. The three-year cumulative multipliers are estimated at 2.28 for GDP, 2 for GVA and 1.17 for private sector GVA. Therefore, given the positive effects of this fund on the economy, I wonder whether an ERDF regional shock could also reduce the financial fragility of the business sector in the Italian provinces during the period of the financial crises (2009-2018). I find that an increase in spending funded by ERDF reduces the non-performing loans to output ratio, especially for the manufacturing and construction sectors. These beneficial effects on credit market risk are greater during periods of credit supply easing and in less competitive regions, especially in terms of innovation.

The third chapter deals with the estimation of multipliers of nationally financed public expenditure in the Italian regions. I use a panel dataset at the NUTS-2 level for the period 1995-2019. I find a cumulative multiplier, over a six-year horizon, of 1.26 which is lower than the estimated GDP multiplier in the previous chapter. I argue that this difference can be attributed to the different expenditure categories considered, i.e. ERDF-financed public investment spending in Chapter 2 and public consumption expenditure in Chapter 3. Indeed, one body of the empirical literature has found higher multipliers generated by public investment rather than public consumption. In addition, I estimate separate multipliers for the Centre-North and the Mezzogiorno. The results give higher multipliers in the most developed area of the country, with a six-year multiplier of about 1.62 in the Centre-North, against a value of 0.94 in the "Mezzogiorno". However, although the difference in multipliers between the two macro-areas is economically important, it is only statistically significant in the short run.

Finally, the last chapter studies the effects of government spending and tax shocks on housing markets and household credit market conditions in Italian regions. The analysis is based on a panel of NUTS-2 level data for the period 2004-2019. I show that government spending expansions and tax cuts improve credit market conditions for households and stimulate the housing market by increasing house prices. Moreover, I find that these positive effects prevail in regions with a higher GDP per capita and in regions with a more developed banking sector. Moreover, fiscal policy plays a counter-cyclical role during and after the GFC, producing higher effects on credit market conditions and house prices in this sample period.

In summary, using data at different territorial levels (NUTS-2 and NUTS-3), observed over different sample periods, I show that regional fiscal policy in Italy improve the credit market conditions with also a broad reduction in the credit risk. The government spending multipliers are found to be positive and higher than one, and even larger when considering the investment expenditure financed by the ERDF fund. Both expenditure and tax policies stimulate the housing markets and relaxes credit market conditions for the households. This set of results has some policy implications, suggesting that, as pointed out by recent empirical studies, fiscal policy may be used as alternative tool to monetary policy, to stimulate the economy and the credit market, especially during period of financial crises, when the monetary policy may be constrained by very low interest rates. However, the results also document that these effects are actually heterogeneous across the country. Fiscal policy is more effective in the more developed area than in the South. This thesis does not study the factors underlying the different economic and structural conditions between the two macro-areas of the country, and I recognise this as a limitation and a starting point for future research, but it does show that these different economic and structural conditions, and thus the different levels of economic and banking sector development that characterise the two macro-areas, influence the way in which fiscal policy impacts on the regions. This requires greater understanding, but also suggests that policymakers should make greater efforts to reduce the secular disparity between the North and the "Mezzogiorno". Furthermore, the multiplier and elasticity estimates should be interpreted as *"geographic cross-sectional fiscal multipliers"*, as they were obtained by pooling regional data and estimating a panel pooled VAR with time and regional fixed effects. Firstly, we used this approach throughout the thesis, following a body of the literature. This allows us to control for unobserved regional heterogeneity and, more importantly, common shocks, such as monetary policy, which is an important advantage when estimating fiscal multipliers. Second, the sample size, in particular the short time series dimension of the panel datasets used, especially in Chapter 4, may limit the use of other tools to account for regional heterogeneity in parameters and to model potential spillover effects between units. However, I acknowledge that neglecting regional heterogeneity in shocks and interrelationships between units is a major limitation of the analysis conducted in this thesis, which deserves to be expanded in the future.

Software and Codes

The results presented in this thesis have been obtained using the software R.

For the first chapter I have used some commands available from the following packages:

- `lpirfs` (Adämmer, 2019) to estimate the IRFs through local projections;
- `plm` (Croissant and Millo, 2008) to estimate panel data models;
- `sandwich` (Zeileis et al. 2020; Zeileis 2006, 2004) and `lmtest` (Zeileis and Hothorn 2002) to compute HAC robust standard errors.

As far as the second chapter, I have written the code for the Proxy-SVAR and for the bootstrap. The local projections estimation is, instead, performed using the above mentioned package `lpirfs` (Adämmer, 2019).

The code for the Proxy-SVAR with the bootstrap, written for the second chapter has also been used for the third chapter, and extended to estimate the dummy-augmented VAR.

The analysis of the fourth chapter has been conducted using codes for the Gibbs sampling and for the sign restriction algorithm that I have written in `R`, based on codes written in `MATLAB` by Blake and Mumtaz (2012), Dieppe et al. (2016) and by lecturers at the Barcelona School of Economics (BSE) summer school on macroeconometrics (Andrea Carriero and Konstantin Boss).

These latter codes are available at the following link:
https://drive.google.com/file/d/1g_IiaOZ7kI8RlntT3aP6YaquZm74TwcQ/view?usp=drive_link.

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CRediT Author Statement

- **Chapter 1** - Francesco Frangiamore: data curation, formal analysis, investigation, software, visualization, writing - original draft. Andrea Cipollini: conceptualization, formal analysis, investigation, methodology, resources, supervision, writing - review & editing.
- **Chapter 2** - Francesco Frangiamore: data curation, formal analysis, investigation, software, methodology, visualization, writing - original draft. Andrea Cipollini: conceptualization, formal analysis, investigation, methodology, resources, supervision, writing - review & editing.
- **Chapter 3** - Francesco Frangiamore: conceptualization, data curation, formal analysis, investigation, software, methodology, visualization, writing - original draft. Marco Maria Matarrese: conceptualization, investigation, resources, visualization, writing - original draft, writing - review & editing. Andrea Cipollini: supervision.
- **Chapter 4** - Francesco Frangiamore: data curation, formal analysis, investigation, software, visualization, writing - original draft. Andrea Cipollini: conceptualization, formal analysis, investigation, methodology, resources, supervision, writing - review & editing.

Publications associated with the thesis

1. Cipollini A, Frangiamore F (2023). Government spending and credit market: Evidence from Italian (NUTS 3) provinces. *Papers in Regional Science, 102*(1), 3-30. (**Chapter 1**).
2. Matarrese MM, Frangiamore F (2023). Italian local fiscal multipliers: Evidence from proxy-SVAR. *Economics Letters, 228*, 111185. (**Chapter 3**).