Does credit crunch investments down?

New evidence on the real effects of the bank-lending channel

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Abstract

We exploit the dramatic liquidity drought in interbank markets following the 2007 financial crisis to identify the effect of credit shocks on firm investments. We focus on a large sample of Italian firms combining information on firm-bank credit relationships with firms and banks balance sheet data. This allows estimating both the direct effect of the liquidity drought on the investment rate and the sensitivity of investment to bank credit. Our findings suggest that the liquidity drought accounts for more than 40% of the negative trend in investments experienced by sampled firms between 2007 and 2010. The effect is stronger for ex-ante constrained firms. We also find that a 10 percentage point fall in credit growth reduces the investment rate by 8-14 points over four years, depending on the definition of the credit variable. Finally, we provide evidence on the effect of credit shocks on firm activity, and on the pass-through to firm's credit chain.

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1 Introduction

The 2007-2008 financial crisis has been followed by the deepest recession since 1930s. Most developed countries featured, in particular, a dramatic drop of investment expenditure: between 2006 and 2010, gross fixed capital formation fell by 10.8% in OECD countries. Because it followed a series of major shocks to banks' liquidity, the drop in investments has been often traced back to a supply-driven contraction of credit (a 'credit crunch'), in that intermediaries proved unable to mitigate the consequences of liquidity shocks on their lending (the so-called "'bank lending channel"', BLC).

Empirically assessing the relevance of the BLC for investments (as opposed to, say, a fall in demand, or deteriorating borrowers' balance-sheets) is a key but challenging task. Identification requires credibly isolating supply from demand determinants of credit growth, and disentangling the role of credit from that of other (observed and unobserved) determinants of investment. And yet, such exercise has a key policy relevance. If the fall in investment is mostly demand-driven, policies that foster private and public consumption would have a direct positive effect on production. The same policies would be much less effective if the drop in investment is mainly due to a credit crunch. Indeed, in this case, policies aimed at improving banks' capital and liquidity position, and at restoring confidence in financial markets, would be more effective in sustaining investments. Moreover, regulatory frameworks aimed at assuring that banks hold sufficient levels of liquidity (such as the Net Stable Funding Ratio criterion imposed by Basel III) may reduce the risk that future financial crises spread over to the real economy.

In this paper, we exploit a large sample of Italian firms (both small and large in size) to estimate the extent to which the BLC is responsible for the fall in firms' investments experienced between 2006 and 2010. While other papers have provided estimate of the effect of the BLC on investments, they have either looked solely at very short-term effects (Almeida et al. 2009, Duchin et al. 2010), or focused only on large and listed firms (Amiti and Weinstein 2013), or provided estimates that may be potentially biased by credit demand effects (Chava and Purnandam 2011, Campello et al. 2010, Gan 2007, Gaiotti 2013).[1]

We focus on the Italian case because of a peculiar feature of the transmission

[1] We defer a thorough discussion of the literature to Section 5.

of the 2007-2008 liquidity crisis to the Italian banking system that allows for a clean identification of the real effects of the BLC. The 2007 and 2008 financial shocks (respectively owed to the subprime mortgage crisis and to Lehman's default) originated outside the Italian economy and hit its banking system through a dramatic liquidity drought in interbank markets. Indeed, as we will argue in Section 2.1, Italian banks were not directly exposed to CDOs, ABS, or Lehman-issued liabilities. As a result, banks that relied more on interbank borrowing *before* the crisis (i.e. at the end of 2006) suffered more from the subsequent liquidity drought (Bonaccorsi and Sette 2013).

We construct a unique dataset that combines detailed information on each firm-bank relationship from the National Credit Register with firm and bank balance-sheet data for a sample of around 30,000 (mostly unlisted) firms. For each firm we measure the pre-crisis exposure to the credit shock as the credit-weighted average of the interbank-to-assets ratio computed in 2006 for all the banks lending to the firm. We provide ample evidence that such a measure of pre-crisis exposure to the financial shock is both significantly related to the subsequent growth rate of credit to the firms, and is not correlated with pre-crisis bank lending strategy, firms' actual and expected investment rates, credit demand, and other observable firm characteristics.

Reassured by such evidence, we perform two main empirical exercises. First, we estimate the direct effect of the BLC (i.e., bank' exposure to the interbank market) on the 2006-2010 firm's investment rate. Second, we use firms' exposure to the credit crunch as an instrumental variable to recover the sensitivity of investment to bank credit. The latter exercise is replicated looking at other firm level outcomes, such as value added, employment, labor cost, expenditures on intermediate output, trade credits and debits, to get a broader picture of the impact of credit supply shocks on the activity of firms.

Our results show a sizeable real effect of the BLC during the crisis: in our preferred estimate a 1 percentage point increase in average interbank-to-assets ratio reduces the investment rate by almost 1 percentage point. We find that this effect is stronger for firms that were ex-ante more likely to be credit constrained (as captured by their cash-holdings, tangible assets, and profitability). As to the estimated sensitivity of the investment rate to bank credit, we find it is highly significant: lowering the growth rate of credit by 10 percentage points reduces the investment rate by 8-14 points, depending on the adopted definition of credit (the lowest bound is obtained restricting to long-term loans, which is probably the most suitable measure of credit when thinking about financing investments). Finally, our estimates suggest that, had the interbank market not collapsed, total aggregate investments by sampled firms in 2007-2010 would have been 47.8% higher.

Our analysis also highlights that the negative credit supply shock induced a significant downsizing of firm's activity, as measured by value added or sales, employment, labor costs and intermediate inputs expenditures. Finally, our results show that a 1 percentage point decrease in bank credit induces firms to reduce trade credit by 0.5 percentage points. The contraction in trade credit is found to be stronger than the one in sales: this finding indicates that the credit crunch may propagate its effect through firms' trade credit chains (as theoretically studied by Kiyotaki and Moore 1997).

Finally, our findings have a direct relevance for the current policy debate about the business model of banks. The exposition of the banking system to the interbank market, while being an effective mean of financing for the economy in normal times, may represent an important source of contagion during financial crises. Thus, the introduction of a Net Stable Funding Ratio, as envisaged in the Basel III regulatory framework, may be a useful precautionary measure to dampen the transmission of shocks from financial markets to the real sector.

The remainder of the paper is structured as follows. Section 2 discusses the empirical strategy implemented to identify the effect of the credit crunch on gross capital formation, and it provides evidence of the validity of the identification hypotheses. Section 3 introduces the data used for the empirical exercises and presents descriptive statistics of it. Section 4 discusses the results, distinguishing between short and medium term effects on investments, heterogeneity analysis, robustness checks, and extensions (the effect of the credit crunch on firm's downsizing and on its credit chain). Section 5 contains a more thorough discussion of the related literature and of the contribution of the paper. Section 6 concludes.

2 Empirical strategy

2.1 The crunch of the interbank market as a source of credit supply shocks

The interbank market is the money market generated by the short-term funding needs of banks, who borrow from banks with excess liquidity. It represents a critical source of funding for intermediaries because it allows banks to readily fill liquidity needs with different maturities (from overnight to more than one year) and through both secured and unsecured contracts. At the end of 2006, total interbank liabilities represented over 13.3% of total assets of Italian banks.[2]

The collapse of the interbank market started on July 2007, when the spread of toxic assets make it impossible for banks to evaluate counterparty risk (Brunnermeier 2009). The situation worsened further after Lehman's default of September 2008. The freeze of the interbank market in Italy can be appreciated by observing the drop in interbank deposits among Italian banks. Figure 2 shows the evolution of total interbank deposits between 1999 and 2010. Total funding in the interbank market topped up at over 24 billions of euros in 2006, while at the end of 2010 it was 4.7 billions: less than 20% of its 2006 value. This fall was driven by increasing counterparty risk, as shown in Figure 3. This Figure plots the trend in the spread between unsecured (Euribor) and secured (Eurepo) interbank lending in Euros. After the Lehman default such spread increased by 3 or 4 times, depending on the maturity, implying a huge increase in the cost of unsecured interbank deposits.

There is large evidence that banks reduce their supply of loans when suffering liquidity shocks, as predicted by the bank lending channel (Kashyap and Stein 2000, Khwaja and Mian 2008). While most papers focused on the case of shocks hitting other sources of funding (e.g. foreign denominated deposits), Iyer et al. (2013) and Bonaccorsi and Sette (2013) show that during the 2007-2008 financial crisis the intensity of the credit supply tightening can be traced to the degree of banks reliance on interbank funding. In fact, the sharp increase in the spread between unsecured and secured interbank transactions signals a widening of the external finance premium for banks. Banks more exposed to unsecured liabilities, such as interbank deposits, suffered relatively more from the increase in the

[2]The distribution of the interbank-to-assets ratio across banks is highly positively skewed, ranging between 0 to more than 80% with a median value of around 1%.

external finance premium.[3] Banks did not prove able at fully substituting interbank funding with other sources, despite policy interventions such as the ECB full allotment auctions (Bank of Italy 2009). The external finance premium paid by banks was then reflected in turn in the cost and availability of funds to borrowers (Bernanke 2007).

This occurred because banks could not fully substitute the fall in interbank transactions with other, cheaper and more readily available, sources. Despite the possibility to access central bank refinancing, in particular after the ECB full allotment auctions (Bank of Italy 2009), interbank markets did not come back to normal functioning, with rates and traded volumes remaining far from their pre-crisis levels especially for longer term transactions, indicating that the external finance premium for banks remained high, despite the ample provision of liquidity by the ECB.[4] Before the financial shock, reliance of banks on interbank funding may have reflected the business model of banks in terms of funding their operations. Nonetheless, in the next Section we will provide evidence that this financing strategies were not correlated with lending strategies before 2007.[5]

Before that, however, notice that Italy is an especially good candidate for identifying the effect of the freeze of the interbank market, because other channels that may confound its identification were largely ignorable. Indeed, Italian banks held little direct exposure towards the assets that became "‘toxic”’ during the crisis (Asset Backed Securities, Collateralized Debt Obligations, etc.), they had little off-balance sheet exposures towards Special Purpose Vehicles, and towards Lehman's liabilities (Bank of Italy 2009). In addition, unlike most other countries, Italy did not experience a real estate bubble (XXX).[6] Hence, Italian banks did not suffer much from losses on mortgages granted to households, households were not hit by adverse wealth effects, and firms were not harmed by reductions in commercial property prices, which may decrease the availability of collateral. In other words, in Italy the shock to bank funding was not amplified by concurrent

[3]Figure 3 indicates that the external finance premium for banks increased across all maturities, even for overnight transactions.

[4]In fact, Brunetti et al. (2013) argue that liquidity provision by central banks crowded out private liquidity, and actually increased uncertainty in markets.

[5]That is, before the crisis, banks were acting consistently with the classical Modigliani-Miller (1958) predictions.

[6]Italian house prices significantly underperformed the boom-bust cycle occurred, for example, in the US, Spain, the UK or Ireland.

shocks on other key asset markets.

2.2 Evidence on the validity of our empirical strategy

Our empirical strategy relies on two important identification assumptions. First, the exposure to the interbank market would have not affected the growth rate of credit during the crisis had the interbank markets not collapsed. Second, before the crisis, exposure was uncorrelated with firms' investment opportunities.

Figure 4 shows the dynamics of the average growth rate of credit (measured with respect to July 2007, i.e. the onset of the crisis) for two groups of banks: those whose interbank-to-assets ratio was above the median of the distribution in 2006, and those below the median. Consistently with our first assumption, the dynamics were very similar until September 2007, when they started diverging: credit from banks with high interbank-to-assets ratio rose at lower pace (and ultimately declined, since January 2009) with respect to credit from less exposed banks.

This graphical evidence can be formally tested using the methodology developed by Khwaja and Mian (2008). We assume that, for each bank-firm relationship, the equilibrium credit flows between year 2006 and any year t can be written as:

$$c_{ij} = \alpha + \beta B_j + d_i + \varepsilon_{ij} \tag{1}$$

where c_{ij} is the growth rate of credit granted to firm i by bank j over the 2006-t time span, B_j is the pre-crisis exposure to the interbank market (measured on December 31^{st} 2006), and the fixed effect d_i captures the change in credit demanded by firm i (and other unobserved characteristics that may affect firm i creditworthiness, as availability of collateral, cash-flow, and the like). The model can only be estimated among those firms that had credit granted from at least two banks at the end of 2006. The identifying assumption is that the demand for bank credit is not 'bank-specific', while banks have differential supply of credit according to firm's observable and unobservable characteristics (Khwaja and Mian 2008).

We estimate model 1 using different time spans. In particular we consider the growth rate of credit granted from 2006 backward till 2002 and forward until 2010. If exposure to interbank market affected credit granted only because of the collapse experienced after mid-2007, we would expect β not to be different from zero before 2007, while it should be significantly negative from that year onwards.

Results are shown in Table 2. Before the crisis (columns 1 to 4) the within-firm growth rate of credit did not differ significantly according to the bank's exposure to the interbank markets. By contrast, from 2007 onwards (columns 5 to 8), credit flows for the same firm started growing at a lower pace the more the bank was reliant on interbank funding. Notice that a Hausman test fails to reject the null of significant differences between a random effect model and the fixed effects one (p-value: 0.37), thus providing evidence that bank's exposure to the interbank markets is not correlated with firm's demand for credit. Nonetheless, in our empirical strategy we will include the estimated fixed effects among the regressors of our investment equations for robustness.

Additional evidence on the role of exposure to interbank market before and during the crisis can be obtained from a firm-level equation:

$$c_i = \bar{\alpha} + \bar{\beta}\bar{B}_i + \gamma d_i + \theta_s + \rho_p + \bar{\varepsilon}_i \tag{2}$$

where c_i is the growth rate of *total* credit granted to firm i, operating in sector s and province p; $\bar{B}_j$ is the weighted average of the initial exposure to the interbank market of banks lending to firm i, with weights equal to their share of credit granted (from now on, "Exposure"); and θ_s and ρ_p are sector and province fixed effects, respectively. The firm-specific demand shock d_i cannot be directly computed in (2). However, an unbiased estimate can be retrieved from equation (1) (Bonaccorsi and Sette 2013).[7]

We single-out two periods: the pre-crisis period, from December 2002 to December 2006, and the crisis period, from December 2006 to December 2010. For each period separately, we estimate the effect of Exposure measured at the beginning of period on the growth rate of credit granted. Consistently with our identification hypothesis, Exposure in 2002 did not affect the growth rate of credit from 2002 to 2006. However, Exposure in 2006 did have a negative and significant effect on the growth rate of credit: a 1 percentage point increase in the interbank-to-assets ratio reduced credit granted by 0.7 percentage points.

[7] An alternative approach is used by Jimenez et al. (2012): they suggest correcting the OLS estimate of $\bar{\beta}$ computed *without* including the firm-specific demand shock in (2) with an estimate of the covariance between B_j and d_i obtained from (1). It is apparent that the two approaches are statistically equivalent: in the remainder of the paper, we follow Bonaccorsi and Sette (2013) and we show among the robustness checks that Jimenez et al. (2012) methodology yields similar results.

Finally, in the robustness checks shown in section 4.4 we will exploit a smaller firm-level dataset (the Bank of Italy's Survey of Industrial and Service Firms - SISF) which include information on firm's investment expectations (measured in April 2007) and we will show that expectations are not significantly correlated with pre-crisis Exposure.

2.3 Empirical model

The availability of bank-firm matched data containing detailed firm balance sheet information allows for a comprehensive assessment of the impact of bank liquidity shocks. First, we augment a standard investment equation by average firm exposure to the shock ($\bar{B}_i$) to obtain a direct estimate of the effect of the bank lending channel on firm's investment (λ):

$$\frac{I_{i,1}}{K_i} = \pi + \lambda \bar{B}_{i,0} + X_i \Phi + \epsilon_{it} \tag{3}$$

where $\frac{I_i}{K_i}$ is the investment rate of firm i over the sample period, and X_i is a matrix of controls which will be detailed below. Importantly, these include the estimated firm-level demand for credit from (1).

We then notice that (3) can be read as the reduced-form expression for a 2-stage approach to estimating the sensitivity of investment to credit growth where firm Exposure is used a source of exogenous variation for bank credit (see section 2.2):

$$\frac{I_{i,1}}{K_{i,0}} = \pi + \delta c_{i,1} + X_i \Phi + \epsilon_{it} \tag{4}$$

where c_i, the average growth rate of credit to firm i over the sample period, is instrumented with $\bar{B}_i$. Conditional on the validity of Exposure as an instrumental variable, δ is an unbiased estimate of the sensitivity of firm investment to bank credit.

Credible estimates of both δ and λ require addressing the well-known estimation issues arising when firm investment opportunities are unobserved. In the empirical investment literature, this problem is largely circumvented relying on (a proxy of) Tobin's Q as a sufficient statistic for investment ratios (see Hayashi 1982). But constructing such proxies implies restricting to listed firms which represent just

1% of firms in our sample. Moreover, Q-based investment regressions have been increasingly subject to several criticisms.[8]

We address the identification issue in several complementary ways. First, we follow recent studies where investment opportunities are captured by low order polynomials in variables (sales, size or measures of profitability) that are available for a larger set of firms. [9] Second, we augment the model with an unbiased estimate of firm's demand for credit: the firm fixed effect d_i estimated in (1). This would capture investment opportunities to the extent that they are correlated with firm's demand for capital. Finally, in section 4.4, we estimate model (4) on the SISF dataset, controlling for self-reported investment opportunities. Results, discussed in Section 4.4, show that our baseline results do not differ.

3 Data

3.1 Datasets

We build our dataset by matching data from three sources. First we obtain balance sheet information of Italian companies, mostly privately held, from the the Company Accounts Data System (CADS). This is a proprietary database, kept by a consortium of Italian banks for credit risk evaluation. The CADS collects detailed balance-sheet information on a large sample of nonfinancial incorporated firms since 1982. The nature of the dataset, routinely used by banks for credit decisions, implies the data are carefully quality controlled. In 2006, firms in CADS accounted for more than 75% of total net revenues by Italian incorporated firms. The sample, however, is not randomly drawn, since a firm enters only by borrowing from one bank.

From CADS we select balance-sheet data from 2006 to 2010 to obtain the main variables we use in our baseline regression (investment, assets, return on assets (ROA)), and other balance-sheet variables that we use in the heterogeneity analysis and for the extensions.

[8] See, among others, Fazzari et al. (1988), Bond and Cummins (2000) and Gilchrist et al. (2005)

[9] Gala and Gomes (2013) show that, under very general assumptions about the nature of technology and markets, the optimal investment policy can be written as a function of low order polynomials in few basic state variables that can be more precisely measured at the firm level, such as sales and size. Alternatively, in line with the early neoclassical literature, Asker et al. (2013) rely on lagged sales growth and on a measure of firm profitability (ROA).

The second source of data is the Italian Credit Register (CR). This is kept by the Bank of Italy (the central bank and banking supervisor) and collects from all intermediaries operating in Italy (banks, other financial intermediaries providing credit, special purpose vehicles) individual data on borrowers with exposures above 75,000 euros towards a single intermediary.[10] The CR contains data on the outstanding bank debt of each borrower, distinguished into loans backed by account receivables, term loans, and revolving credit lines. The CR also contains information about the granting institution and the unique tax identification number of the borrower. The quality of the CR data is ensured by the fact that banks routinely use the CR as a tool to monitor borrowers. We select all credit relationships between banks and firms in each year from 2006 to 2010. We also select data back to 2002 to run the placebo regressions.

The third source of data is the Supervisory Reports submitted by banks to the Bank of Italy. These contain balance-sheet data of all banks operating in Italy, including banks that are not listed on the stock market. From these data we select interbank deposits taken by each bank and total bank assets at December 2006 (at December 2005 for the placebo regression), to construct the interbank to assets ratio, on which we base our instrument for credit growth. We use consolidated data, to exclude interbank deposits made to banks belonging to the same banking group.

3.2 Sample selection

First, we match data on each bank-firm relationship from the CR with data on banks' interbank to asset ratio from the supervisor report using the unique bank identification number ('ABI code'). Then we aggregate data on all loans to each firm from the CR and we match them with firm balance sheet data using firms' unique tax identification number.

We exclude subsidiaries of foreign banks since they fund their activity almost exclusively through interbank transactions from the headquarter, and we cannot distinguish true external interbank funding from internal transfer of funds. Subsidiaries of foreign banks grant only a small share (about 6 percent) of total loans

[10]Exposures include both debt and guarantees. A borrower with debt of, say, 20,000 euros towards a bank appears in the CR if she also provides guarantees worth at least 55,000 euros to another individual borrowing from the same bank. The 75,000 euros threshold has been decreased to 30,000 since January 2009.

to Italian firms.

To estimate fixed effects in model (1), we restrict our attention to firms that obtain loans from at least two banks, as in, among others, Khwaja and Mian (2008) and Jimenez et al. (2012). Multiple banking is common in Italy, even among small firms (Detragiache et al. (2000), Gobbi and Sette (2013)).[11] Focussing on firms borrowing from at least two banks goes against finding an effect of credit on investment since single bank firms are more likely to be credit constrained. After this step we are left with XXX firms.

Finally, we include firms that are active in all years from 2006 to 2010, to compute investment over the crisis period. This amounts to excluding firms that disappear from the sample. If the probability of exiting the market is higher for credit constrained firms, our estimates are a lower bound of the full effect of credit availability on investment. Overall, the sample we use in the baseline regression includes 29,132 firms.[12]

Table 1 shows descriptive statistics of the firms included in the sample. The 2007-2010 investment rate is equal to cumulative investments over the 4 years normalized by firm assets.[13] Data indicate that the median firm makes gross investment equal to its initial capital in the 4 years period. This is reasonable, since it implies an average yearly gross investment equal to about 25 percent of fixed assets, which is in line with the evidence from the US and Japan.[14]

Credit growth is the cumulative growth rate of credit granted (commitments) from December 2006 to December 2010. [15] On average, credit granted grew over the period, although almost half of the firms experienced a contraction in credit granted.[16]

[11] Gobbi and Sette (2013) using a similar sample, find that about 7 percent of the firms in the CADS have only one banking relationships. Such firms are smaller, less leveraged, invest a smaller fraction of their revenues than the average firm.

[12] The inclusion of all firm-level controls additionally excludes 9 firms from the sample, as they did not report complete balance sheet data in CADS.

[13] We trim the top 10% observations, as the book value of capital and investment are extremely noisy, and we want to avoid our results to be driven by outliers. We test the robustness of all results to different trimming thresholds, and to winsorize data instead of trimming, in Section 4.

[14] The investment to asset (book capital) ratio of large Japanese firms in Gan (2007b) is on average 31 percent; that of US Compustat firms used in Almeida and Campello (2007) is around 25-30 percent.

[15] In computing the growth rate of credit, we keep track of existing credit relationships over time even after a bank disappears from the sample due to a merger or an acquisition. In this case, we assume the firm had a relationship with the new bank from the beginning.

[16] Since in our data yearly growth rates of credit granted higher than 10% are widespread, in the baseline specification we compute the actual growth rate for each bank-firm relationship that is present

All other variables refer to 2006 (end of year data). In particular, the average exposure to the interbank market (interbank funding to bank assets ratio) is around 12 percent. Firms are small (median fixed assets are 2.1 million euros, about 2.7 million US Dollars), and the vast majority of them are not listed (only X percent are listed on the stock market).

The table also shows the distribution of sampled firms by industry: manufacturing represents more than half of the sample, services about 40, construction about 7 per cent.

Finally, the table shows the distribution of other outcome variables that may be affected by the credit crunch: value added, employment (average number of employees during the year), labor costs, purchase of intermediate inputs, and trade debit and credit. Notice that information on the number of employees is available only for a subsample of 17,486 firms. We test the effect of the credit supply shock on these variables in Section 4.5.

4 Results

4.1 Baseline results

We first estimate the baseline model (3). Results are shown in Table 3. Exposure has a negative and significant effect on investment in all specifications. Column 1 shows estimates from the baseline model without controls. A 1 percentage point increase in Exposure reduces the four-years investment rate by around 1 percentage point. The firm fixed effect retrieved from equation (1) has a positive and significant coefficient, consistent with it capturing firm-level demand for credit.

The coefficient of Exposure remains practically unchanged when firm-level controls are included in the regression (columns 2 to 6), providing support to the hypothesis that Exposure is not correlated with firm characteristics. In Column 2 we include a first set of firm-level controls. Following Gala and Gomes (2013) we include fixed assets (linear and squared to account for potential non-linearities of the effect of size), the sales to assets ratio, and the investment ratio in 2006, to proxy for investment opportunities. The coefficient of Exposure is still negative

before the crisis. Among the robustness checks, however, we will test that our result are robust even to a log-differences specification, which is a commonly used approximation (see, among others, Khwaja and Mian 2008 and Jimenez et al. 2012).

and significant at the 5% level: a 1% increase induces a decrease of the investment-rate by 0.9% over four years. Column 3 also includes the cash holdings to assets ratio, a commonly used control in investment equations that accounts for credit constraints. The estimate in Column 4 controls for the growth rate of sales, as in a standard investment accelerator-model (Bernanke et al. 1999), while Column 5 adds cash holdings to account for imperfect access to capital markets. Finally, Column 6, our preferred estimate, proxies Tobin's Q with firm's Returns-On-Assets (ROA), as in Asker et al. (2013).

The estimated effect of Exposure is not only statistically significant but also economically relevant. Based on the coefficient in Column 7, for example, we estimate that the drop in total investments by the sampled firms induced by the credit crunch amounts 48.4% of total investment expenditure between 2007 and 2010. [17].

Finally, the last row of Table 3 provides, for each model, the standardized effect of average interbank exposure on firm's investment rate. The estimated effect remains stable over the different specifications: a 1 standard deviation increase in Exposure induces a 0.2% standard deviation reduction in the investment rate.

4.2 Heterogeneity

The average effect estimated in Table 3 could be the result of heterogeneous reactions to the credit crunch.

We first distinguish firms on the basis of the industry they belong to. To study this dimension of heterogeneity, we interact Exposure with industry dummies. Column 1 of Table 4 reports the corresponding point estimates, together with standardized effects to ease the comparison between results from different subsamples. The effect of Exposure is stronger for manufacturing firms and for services firms. It is weaker for construction, likely because these firms have a higher availability of collateral.[18]

Other important dimensions of heterogeneity are the ex-ante liquidity of firms, the pledgeability of assets, and firm profitability. To proxy for these characteristics we use, respectively, cash-holdings over fixed assets, tangible over total fixed assets,

[17]This is obtained as follows: we first multiply the predicted drop in investments due to the average interbank exposure for each firm ($\delta \times \bar{B}_{i,0}$) by the firm's pre-crisis assets $K_{i,0}$ to obtain the predicted drop in investments for each firm. These predictions are then aggregated over the sample.

[18]It should be recalled that Italy did not experience a housing bubble (see Section 2.1).

and EBITDA over value added, all measured at the end of 2006. We first regress each of these variables on the baseline set of pre-crisis controls (assets, ROA, sales, and investment rate), to purge the heterogeneity analysis from other possible confounding factors that affect their distribution.[19] Then, we distinguish firms that are below and above the median of this (conditional) distribution, and we interact Exposure on a dummy that identifies these two groups to estimate heterogeneous effects. Results, provided in Columns 2 to 4 of Table 4, indicate that Exposure has a significant effect on investment only for firms with below-median cash-holdings, tangible assets, and EBITDA. This is consistent with the idea that such firms are less able to substitute bank credit with other sources of finance, because they lack substantial buffers of cash, because they have less pledgeable collateral, because they generate lower cash-flows.

Finally, as an additional exercise, we estimate heterogeneous effects in the short and the medium term using as a dependent variable investment rates estimated over different time-windows.[20] The estimated effects for below-median and above-median cash-holdings are plotted in Figure 5. The effect for low cash-holding firms increases over time, it is significantly different from zero from 2009 onwards (in 2008 it has $p < 0.10$). Conversely, high cash-holding firms do not experiment any drop in investments induced by the credit crunch, neither in the short nor in the medium term.

4.3 IV: the sensitivity of investment to bank credit

We now turn to estimating the sensitivity of the investment rate to the growth rate of bank credit. Credit at the firm-level is likely to be endogenous, and simple OLS estimates would be biased. However, the direction of the bias is ex-ante ambiguous. OLS may be upwardly biased if higher investment rates induce higher demand for credit (reverse causality), or if banks prefer to lend to more profitable firms, and this expected profitability is positively correlated with investments (omitted variable bias). At the other side, a negative economic downturn may reduce both investment opportunities and cash-flows, and firms may use more intensively external finance to sustain working capital. In this case, firms may contemporaneously

[19] Heterogeneity analyses based on unconditional distribution of cash-holdings, tangible assets, and EBITDA are qualitatively similar, though less precisely estimated.

[20] The firm fixed effect has been estimated differently for each year, too.

cut investment and increase their demand for bank credit, thus inducing a downward bias in OLS estimates. This downward bias has been detected by the trade literature studying the effects of credit on export (see e.g. Paravisini et al., 2013), Federico, 2013).[21]

To address the endogeneity of credit growth, we use Exposure as an instrument for credit growth, and estimate model (4) via 2SLS. We include the same controls as the baseline regression. Results are shown in Table 5. Column 1 reports OLS correlation: a 1% increase in credit growth is associated with a 0.5% increase in the four-years investment rate. However, IV estimate of the sensitivity of investment to bank credit, provided in Column 2, are almost three times higher. A 1% increase in credit growth raises investment rate by 1.4%. The impact is sizeable even in terms of standardized effect: one standard deviation increase in credit growth raises the investment rate by 37.6% of a standad deviation. Thus, our findings are consistent with a downward bias in OLS estimate. Notice that Exposure is a strong instrument for credit growth, as its F-statistics in the first stage is fairly high (43.97).

The relatively large size of the IV estimates is not surprising, given that: (i) capital expenditure is heavily dependent on bank credit (i.e. other sources of finance are not easily available), and (ii) bank credit is used to finance other than capital expenditures (e.g. working capital) which might be more difficult (i.e. more rigid) to cut following a credit shortage. This latter point suggest that the sensitivity of investment rates to credit growth would be lower should we be able to more precisely measure the specific component of bank credit that firms use to finance investments. To this purpose we restrict our attention to the long term component of total credit (i.e. term loans). Term loans include mortgages and leasing, have longer maturities and their dynamics are therefore more directly linked to investment decisions. Results obtained using the growth rate of term loans as our endogenous variable are shown in Column 3 of Table 5. The estimated sensitivity is still positive and highly significant, but drops to around 0.7. A 1 standard deviation increase in the growth rate of long-term credit leads to an increase of 43.1% of a standard deviation in the investment rate.

Interestingly, our estimates are larger than those obtained by Amiti and We-

[21] Paravisini et al. (2013) note that "a collapse in the prices and demand for a firm's exports reduces substantially the cash-flows generated by the firm internally through revenues. To substitute for this decline in internally generated cash, firm's demand for external finance increases".

instein (2013) for the case of Japan.[22] These differences can be due to several factors. First, they focus on a sample of listed firms only, while we look at a large pool of mostly small-medium size firms, for which bank credit is more relevant.[23] . Second, access to other sources of finance is likely more developed in Japan than in Italy, since capital markets are more developed, and firms are larger in the former than in the latter. Thinking of both the development of capital markets and the firm size distribution across countries, this reasoning also suggests that the real effects of a credit supply shock will be larger in the case of other large European countries, such as Spain or France, and smaller in the US, or the UK. Hence, our results might be ultimately helpful in understanding the different real consequences of the crisis in different areas of the world.

4.4 Robustness checks

We test the robustness of our results to several checks. In Panel A of Table 6, we change the size and composition of the sample.[24] As discussed in Section 3, our dependent variable is trimmed at 10% to reduce the influence of outliers. If we trim a smaller share of observations, outliers affect markedly all estimated moments of the distribution. By trimming the top 5%, for instance, the average investment rate reaches 300.9, and its variance tops up at 874.4: three times the mean and variance of the 10% trimmed distribution. The first Column of Panel A in Table 6 shows results of estimating our baseline model on the 5% trimmed sample. The point estimate is significant and five times higher with respect to the one obtained in the 10% trimmed sample. This increase, however, balances the higher variance of the dependent variable. Thus, the resulting standardized effect is similar to the one obtained in our baseline results of Table 3: a 1 standard deviation increase in Exposure raises the investment rate by 0.2% of a standard deviation. Similar results are obtained if we winsorize the distribution, instead of trimming it, at 5%

[22]Comparison of their results with ours is complicated by the fact that Amiti and Weinstein (2013) only report the sensitivity of investment to credit supply shocks interacted with other firm-level variables, such as loans-to-assets ratio. However, descriptive statistics of loans-to-assets ratio are provided in the paper. The sensitivity to credit supply shocks measured at the mean of the interacting variable is a tiny 0.049, while the one measured at the maximum of it is 0.495. In short, we can safely conclude that our estimated sensitivity is larger than theirs.

[23]Consistently, Asker et al (2013) found that listed firms are less responsive to changes in investment opportunities, even during the recent financial crisis.

[24]Unless otherwise indicated, all the subsequent estimate include the controls in Column 6 of Table 3. However, results are robust to all the baseline specifications.

or 10% (Columns 2 and 3, respectively).
Some 338 firms were excluded from the sample because they did not provide information on investments for some years between 2007 and 2010. In the fourth Column of Panel A we include them and use the average yearly investment rate, instead of the cumulative rate, as our dependent variable. A 1% increase in Exposure reduces yearly investment rate by 0.27%, which corresponds to the 1.1% compound effect estimated in Table 3.

Panel B of Table 6 tests the robustness of our results to changes in the baseline model. In Column 1 we include sector-times-province fixed effect. This specification may be more robust with respect our baseline because firms belonging to the same sector and located in the same province are likely to face similar demand shocks.[25] Though less precisely estimated (p-value = 0.06) the coefficient remains remarkably similar to the ones of Table 3, in terms of both point estimate and standardized effect.

Column 2 of Panel B shows results obtained by weighting observations by firm sales; that is, giving larger weights to larger firms in the estimate. This may be important as the size distribution of firms in our sample is positively skewed, and the results might therefore be driven by small firms only. Nonetheless, results are robust to this weighting: the coefficient of exposure is still negative and significant (now at the 1% level) and the standardized effect is larger: one standard deviation increase in interbank exposure induces an increase of 4.5% of a standard deviation in the (sales-weighted) investment rate.

A potential issue with our identification strategy is the presence of non-random sorting between banks and firms that may be correlated with Exposure and firm's investment opportunities. This may happen if, for instance, larger banks would be more exposed to interbank markets and they would be better able to estimate investment opportunities. In this case, our estimate would be biased downward. In the third column of Panel B, we include to our baseline specification a set of fixed effects for the main bank lending to the firm. Main bank is defined as the bank that was lending the highest share of bank credit at the end of 2006. Though this only partially controls for this identification issue, we would expect that, if sorting is a relevant phenomenon, the estimated coefficient would change by the

[25] In principle, the additive structure of the province and sector fixed effects in our baseline model may upwardly bias the estimated effect if, for instance, banks that were less exposed to the interbank market were more able to identify sector profitability at the province level.

inclusion of this fixed effects. However, the point estimate remains similar to that of the baseline and statistically significant.

So far, we have measured Exposure as the ratio between average gross interbank liabilities and firm's total assets. In Column 4 of Panel B, we consider *net* interbank liabilities, normalized by firm's assets. This may be important since, in principle banks may simultaneously borrow and lend on the interbank market. Using the net interbank position does not affect results.

Finally, a bias may emerge if firms pre-crisis investment opportunities are correlated with interbank-to-assets ratio of the banks lending to it. This may happen if, for instance, larger investment opportunities by firms exercise pressure on bank funding. Although in Section 2.1 we have shown that our Exposure is not correlated with firms demand for credit (as proxied by firm fixed effect), we can provide more direct evidence that exposure is not correlated with expected investment opportunities using the Bank of Italy Survey of Investment of Industrial Firms (INVIND). INVIND is an annual representative survey of medium and large Italian firms from manufacturing, construction, and private services sectors. It collects information both on the actual level of investments of the past year and on its expected levels at the end of the present year. Crucially for our analysis, it is administered in April of every year: i.e., for 2007, *before* the onset of the global financial crisis. Hence, we can estimate models (3) and (4) using as a dependent variable the expected growth rate of investments from 2006 to 2007. Results are shown in Table 7. In Column 1, we regress the expected growth rate of investment on Exposure and show that they are not significantly correlated. Second, we replicate our baseline model with INVIND data. Because of imperfect matching between the Credit Register and INVIND, and of panel attrition over the 2006-2010 period, the final dataset is composed of 996 firms (around 1/3 of the cross-sectional sample). The result (Column 2) is consistent with our baseline finding: a 1% increase in Exposure reduces the investment rate by 0.8%. The inclusion of investment expectations among the controls (Column 3) increases the precision of the estimate, but does not change significantly the point estimate. Finally, Column 4 performs the 2SLS model to identify the sensitivity of investments to a credit supply shock. The resulting point estimate is less precisely estimated than the one obtained from the full sample. Though the point estimate is smaller, the standardized effect is very similar to the one obtained in Section 4.3: a 1 standard deviation increase in bank credit raises the investment rate by 49% of a standard deviation.

4.5 Extensions

We extend our main analysis to test whether the credit supply shock had an effect on other outcome variables than investment. Indeed, as discussed in Section 2.3, a prolonged drop in bank credit may determine a significant downsizing of the firm. To look for evidence of this effect, we consider balance-sheet data on labor cost, expenditures on intermediate inputs, value added, and sales. In addition, for a subsample of 17,486 firms, we obtained data on average yearly employment for the 2006-2010 period. We check whether the credit supply shock led to a drop in these variables.[26] Results, based on model (4), are shown in Table 8. Credit growth significantly affects firm's value added: a 1% increase in total credit granted raises value added by 0.28%.[27] This sensitivity is backed by all other outcome variables: credit growth has a positive and significant effect on employment, labor costs, and intermediate expenditures (Columns 2-4)

This reduction in the factors of production is backed by a significant reduction in final production as measured by value-added or sales. Credit growth significantly affects the total amount of firms expenditures on wages and on purchases of intermediate goods and services. Indeed, a 1 percentage point lower credit growth causes a reduction of 0.16% in total wages, and of 0.11% of expenditure in intermediate goods and services. Thus, the credit crunch seems to have had an effect also on the overall scale of firm activity.

We then look at the activity of firms as providers of trade credit to customers and receivers of trade credit from suppliers, as part of a credit chain (Kiyotaki and Moore 1997).[28] This may represent an important amplification mechanism of the shock: a crunch in bank credit may be reflected into a drop of trade credit. Results, shown in Columns 5 and 6 of Table 8, indicate that firms borrowing from banks more affected by the crisis granted less trade credit to their customers: a 1 percentage point decrease in the growth rate of bank credit to the firms reduced

[26] We even replicated all estimate in the subsample of observations for which employment is non-missing. Results, available upon request, are qualitatively and quantitatively similar to those provided here.

[27] The corresponding estimate for total sales (available upon request) is 0.15%.

[28] See also Garcia-Appendini and Montoriol-Garriga (2012) for evidence on the effect of the 2007-2008 crisis on trade credit.

the growth rate of commercial lending by the firm by 0.5 percentage points. Conversely, we do not find any significant effect on the amount of trade debit that firm receive.

The negative effect of the credit crunch on trade credit may be partially explained by firms downsizing: if firms reduce their sales, then the volume of credits to their customers may reduce proportionally. To test whether this is the case, we use as the dependent variable the difference between the growth rate of trade credits and the growth rate of sales. Results, provided in Column 7, show that this is not the case: a positive (negative) credit shock raises (reduces) trade credits more than sales.

5 Discussion and related literature

Our research contributes to different strands of the literature.

First, the one on the real effects of financial crises. Most of this literature relies on a sample-split strategy: first, firms that were ex-ante credit constrained are identified (either through self-assessment (as in Campello et al. 2010 and Gaiotti (2013)) or through proxies such as cash-flow and liquidity measures (as in Duchin et al. 2010, among others)); second, the performance of firms identified as credit constrained is compared to that of unconstrained ones during the crisis.[29] We extend these works in two dimensions. First, we identify a proxy for exposure to the credit crunch that is not correlated with firms' characteristics. Indeed, while during periods of positive economic growth the general improvement in borrowers' and lenders' balance sheets may blur the difference between credit constrained and unconstrained firms, during crises the latter may outperform the former simply because the fall in demand affects more the worse-off firms.

An approach more similar to ours is pursued by Gan (2007a) and (2007b), using the impact of the early 1990s burst of a real estate bubble in Japan to study the impact of a credit shock on investment. Gan finds that firms borrowing from banks more exposed to the real estate market invested less. While our methodology is conceptually similar to hers, we extend her results in three ways: first, we can explicitly control for firm's credit demand; second, we derive the sensitivity of of

[29]Usually, these empirical exercises are corroborated by testing whether the two groups of firms displayed the same trend before the financial shock.

investment to credit, and not only the reduced form estimate of the exposure to the crisis on investment. Third, and more importantly, the Japanese real estate bubble used by Gan may be affecting directly both bank's liquidity *and* firms' collateral, in this way mixing the bank lending channel with the collateral channel (Chaney et al. 2010).

Finally, a very recent paper by Amiti and Weinstein (2013) has developed a methodology for decomposing credit flows in supply and demand shocks (identifying both their idiosyncratic and collective components). They estimate it using a small sample of listed Japanese firms. As contribution with respect to this paper (and to most of the previously cited, aswell), we provide the first evidence of the impact of the credit crunch on a sample of small and medium sized firms, mostly unlisted, that are highly dependent on bank credit for external finance. While small and medium sized enterprises play a key role even in market oriented financial systems (Berger and Udell 1998), there is little evidence on their investment behavior, in particular during crisis, mostly because of lack of data.

Our paper also contributes to the large literature on the effect of credit constraints on investment. From the seminal paper of Fazzari et al. (1988), several papers proposed alternative proxies for credit constraints, and ways to address the endogeneity of the measures of credit constraints. A key tenet of the approach of Fazzari et al. is that the sensitivity of investment to internal finance is larger for more credit constrained firms. Kaplan and Zingales (1997) argue that this monotonicty is not a necessary feature of investment by credit constrained firms. Several subsequent work (among others Lamont (1997), and Almeida and Campello (2007)) attempt to make progress to find measures of credit constraints. We contribute to this literature by using a direct measure of access to credit and by improving on the identification of the effect of credit on investment, thanks both to our data, and to the use of an exogenous shock to bank lending.

Finally, our work is also related to the large, and growing, literature on the bank lending channel. Our identification strategy is similar to the technique pioneered by Khwaja and Mian (2008), though it is applied to a between-firms comparison. The presence of a bank lending channel has been first documented by Bernanke et al. (1993) and Kahyap and Stein (2000). Progress on identifying supply from demand effects has been made by Peek and Rosengren (2005) who use a sharp decline in Japanese stock prices as an exogenous shock to the supply of credit by

Japanese-owned banks in the US. They find that US branches of Japanese banks cut credit when their parent banks experienced losses in their holdings of stocks, tightening their capital requirements. Further results on the importance of the bank lending channel have been provided by Jimenez et al. (2012), Iyer et al. (2012), Bonaccorsi and Sette (2012), using credit register data and identification strategies à la Khwaja and Mian (2008). These papers find significant supply effects of shocks to the cost of bank funding due either to monetary policy, or to higher interbank rates during the 2007-2009 financial crisis. We extend these results by looking at the real effects of the bank lending channel.

6 Conclusions

In this paper we have provided new evidence on the effect of the 2007-2008 credit crunch on firm investment rate over the 2007-2010 period. We focus on Italy, a country where the financial crisis hit the banking system mainly through a significant liquidity drought in interbank markets. We use a large sample of Italian firms, for which we observe both bank-firm relationships and balance sheets data before and during the current economic downturn. For each firm in our sample, we proxy the exposure to a negative credit supply shock during the crisis with the pre-crisis average exposition to the interbank markets of banks lending to firm. Average exposure to interbank markets significantly predicts both a drop in credit granted and a sizeable fall in the investment rate. The latter effect is stronger for firms that were ex-ante more likely to be credit constrained, as measured by pre-crisis cash-holdings over assets, tangible-to-total assets, and EBITDA-to-value added ratios.

We then use interbank exposition to instrument the growth rate of credit granted, and estimate the sensitivity of investments to bank credit.

Finally, we provide evidence of firm downsizing over the medium term in response to the credit crunch, and evidence on the impact of this crunch on firms credit chain.

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8 Figures and Tables

Figure 1: Growth of credit to the private sector and of GDP in Italy

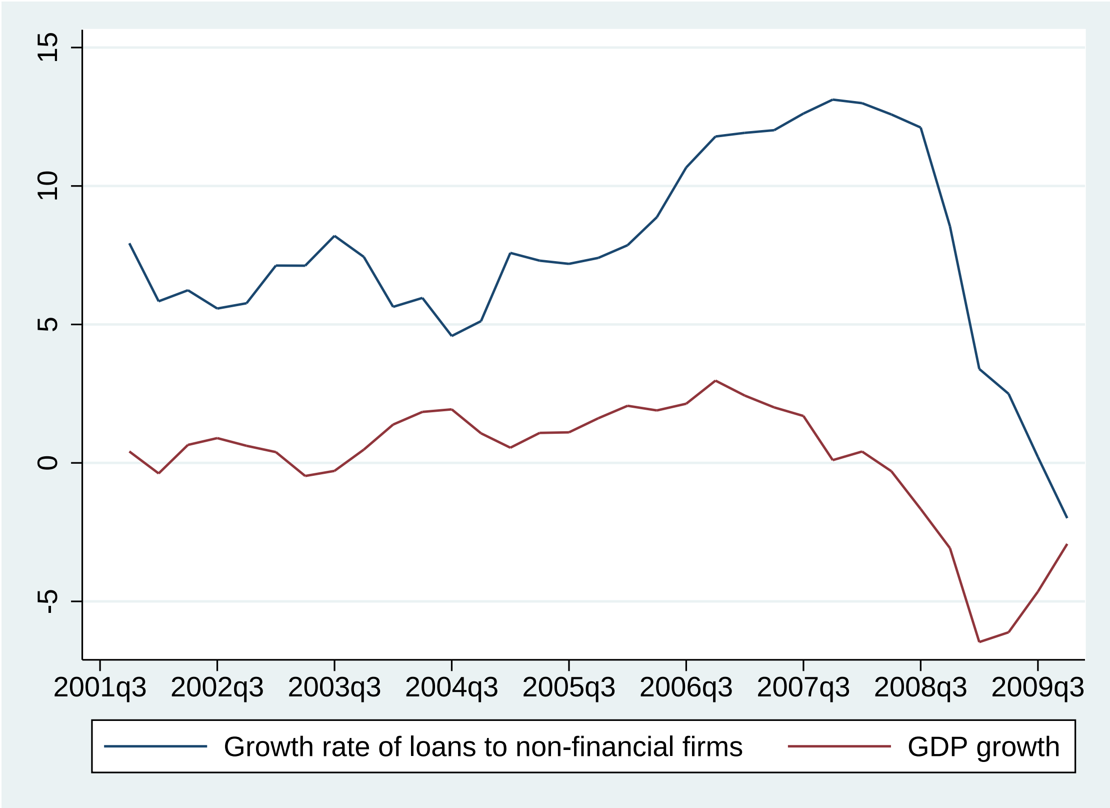

Figure 2: Total interbank deposits of Italian banks (e-MID market) in constant 2005 euros

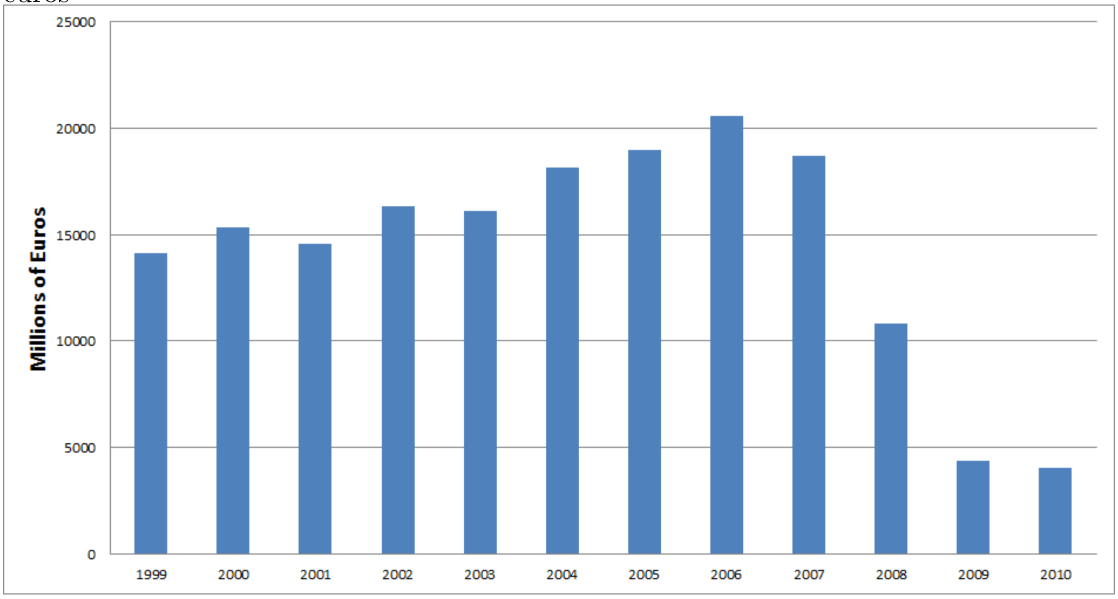

Figure 3: Euribor-Eurepo spread

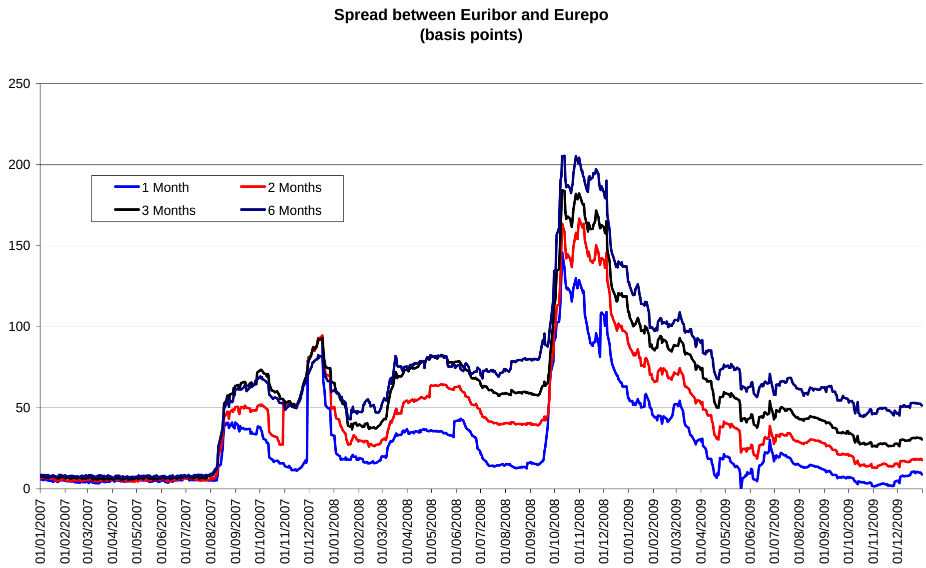

Figure 4: Growth of credit granted by banks above and below median interbank/asset ratio

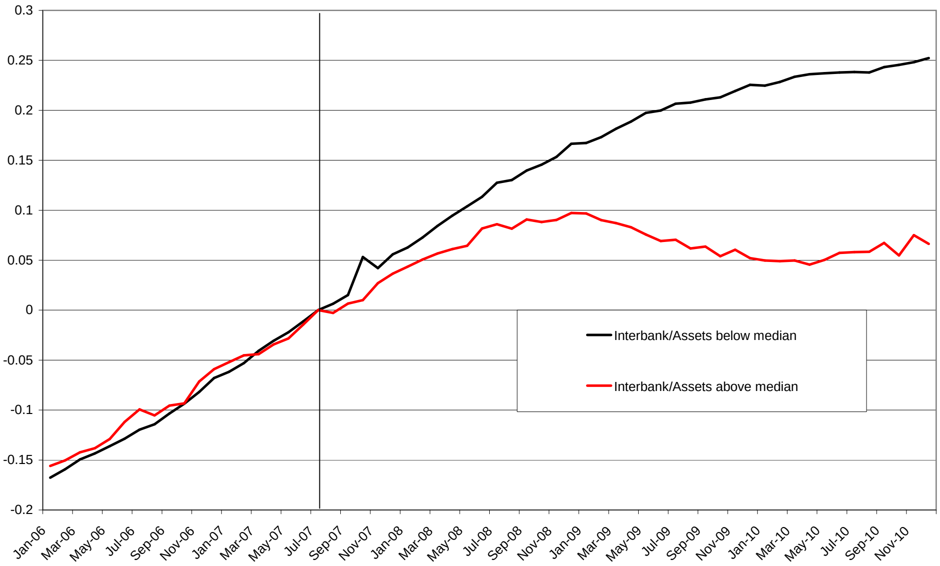

Figure 5: Effect of average exposure to the interbank market on growth rate of gross capital overtime - for firms with below-media and above-median cash-holdings-to-assets ratio (solid and dashed lines, respectively).

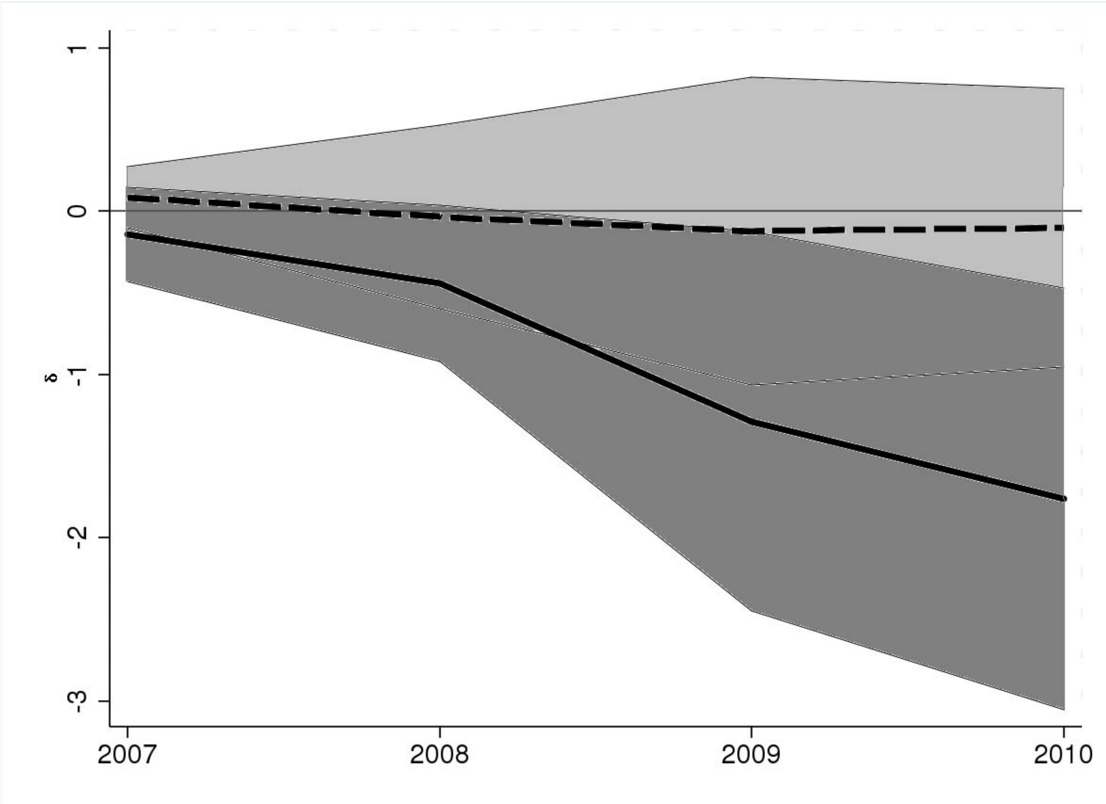

Table 1: Descriptive Statistics

	Mean	St.Dev.	Min	Max	Obs.
Investment Rate (%)	98.30	199.63	-99.40	1382.13	29123
Credit Growth (%)	6.78	51.61	-100	180.68	28227
Avg. Exposure to the Interbank Market (%)	11.83	3.41	.003	55.63	29123
Fixed Assets (000)	24985.63	804586.70	0	105902600	29123
ROA	6.15	7.05	-26.59	40.28	29123
Cash Holdings/Assets	1.57	12.27	0	305.09	29123
Sales/Assets	40.80	265.12	0	5881	29123
Investment Rate 2006	163.93	1019.25	0	12225	29123
Construction	.07	.26	0	1	28693
Tertiary	.39	.49	0	1	28693
Material/Assets	.81	.32	0	1	29123
EBITDA/VA	.40	2.76	-139	240.06	29113
Growth Rate of Long-Term Credit (%)	26.32	108.72	-100	506.52	25414
Growth Rate of Value Added (%)	3.39	43.55	-99.99	135.12	
Growth Rate of Employees (%)	-2.82	23.13	-99.23	46.06	
Growth Rate of Labor Cost (%)	10.16	34.64	-99.98	105.28	27193
Growth Rate of Intermediate Exp. (%)	-6.99	34.51	-99.92	81.32	27009
Growth Rate of Sales (%)	-6.59	35.83	-1553.12	77.91	27530
Growth Rate of Trade Credits (%)	2.87	46.45	-99.99	146.03	25709
Growth Rate of Trade Debits (%)	-2.95	44.45	-99.96	129.31	25634

Table 2: Effect of Banks' Exposure to the Interbank Market on the Growth Rate of Credit Granted, Before and During the Crisis

	(1) 2002	(2) 2003	(3) 2004	(4) 2005	(5) 2007	(6) 2008	(7) 2009	(8) 2010	(9) 2002-2006	(10) 2006-2010
Bank Exposure in 2006 (bank level)	-0.130	-0.127	-0.056	0.055	-0.309**	-0.625***	-0.295***	-0.679***		
	(0.240)	(0.219)	(0.182)	(0.089)	(0.131)	(0.163)	(0.101)	(0.121)		
Avg. Bank Exposure in 2002 (firm level)									-0.122	
									(0.130)	
Avg. Bank Exposure in 2006 (firm level)										-0.703***
										(0.123)
Firm Fixed-Effect	Y	Y	Y	Y	Y	Y	Y	Y	0.913***	0.880***
									(0.012)	(0.011)
Sector FE	N	N	N	N	N	N	N	N	Y	Y
Province FE	N	N	N	N	N	N	N	N	Y	Y
No. of Obs.	149972	161892	173500	187570	230189	223537	221283	221525	40964	31602

Notes: The dependent variable in columns 1-8 is the credit growth between each year and year 2006. The dependent variable in columns 9-10 are the credit growth in period 2002-2006 and 2006-2010, respectively. Heteroskedasticity robust standard errors clustered at the bank level (columns 1-8) and at the main bank and sector levels (column 9-10) in parentheses with *** p<0.01, ** p<0.05, * p<0.1.

Table 3: Effect of Banks' Exposure to the Interbank Market on the 2007-2010 Investment Rate

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Exposure	-1.054**	-0.919**	-0.909**	-1.173***	-1.120***	-0.998**	
	(0.485)	(0.428)	(0.449)	(0.436)	(0.421)	(0.432)	
	[0.487]	[0.431]	[0.452]	[0.439]	[0.423]	[0.434]	
Firm Fixed-Effect	0.355***	0.383***	0.384***	0.331***	0.362***	0.364***	
	(0.036)	(0.031)	(0.030)	(0.035)	(0.030)	(0.030)	
	[0.080]	[0.072]	[0.072]	[0.080]	[0.073]	[0.073]	
Fixed Assets		-0.017**	-0.024***			-0.023***	
		(0.007)	(0.008)			(0.008)	
		[0.007]	[0.008]			[0.008]	
Fixed Assets Sq.		0.000**	0.000***			0.000***	
		(0.000)	(0.000)			(0.000)	
		[0.000]	[0.000]			[0.000]	
Sales/Assets		0.696***	0.105***			0.106***	
		(0.096)	(0.023)			(0.023)	
		[0.096]	[0.023]			[0.023]	
Sales/Assets Sq.		-0.000***					
		(0.000)					
		[0.000]					
Invest Rate 2006		0.008***	0.010***			0.010***	
		(0.001)	(0.001)			(0.001)	
		[0.001]	[0.001]			[0.001]	
Cash Holds/Assets			1.416***		2.965***	1.391***	
			(0.252)		(0.412)	(0.246)	
			[0.253]		[0.413]	[0.247]	
Sales Growth Rate				-0.000	-0.000		
				(0.000)	(0.000)		
				[0.000]	[0.000]		
ROA				1.444***	1.347***	1.359***	
				(0.156)	(0.169)	(0.175)	
				[0.157]	[0.170]	[0.176]	
Sector FE	Y	Y	Y	Y	Y	Y	Y
Province FE	Y	Y	Y	Y	Y	Y	
No. of Obs.	29132	29123	29123	28946	28946	29123	
Std. Effect	-0.018	-0.016	-0.016	-0.016	-0.021	-0.020	-0.018

Notes: The dependent variable is the gross growth rate of capital between 2006 and 2010. Heteroskedasticity robust standard errors clustered at the main bank and sector levels in parentheses with *** p<0.01, ** p<0.05, * p<0.1. Bootstrapped standard errors based on 200 replications in brackets.

Table 4: Heterogeneity in the Effect of Banks' Exposure to the Interbank Market on the 2007-2010 Investment Rate

	(1) by Sector	(2) by Cash Holdings/Assets	(3) by Material As./Assets	(4) by EBITDA/VA
Manufacturing	-1.167*			
	(0.639)			
Construction	-0.365***			
	(0.067)			
Tertiary	-0.961**			
	(0.447)			
Below Median		-1.779**	-1.420**	-1.544***
		(0.696)	(0.663)	(0.405)
Above Median		-0.147	-0.298	-0.561
		(0.454)	(0.445)	(0.553)
Controls	Y	Y	Y	Y
Sector FE	Y	Y	Y	Y
Province FE	Y	Y	Y	Y
No. of Obs.	28682	29112	29112	29102
Std. Effect - Manufacturing	-0.020			
Std. Effect - Construction	-0.006			
Std. Effect - Tertiary	-0.016			
Std. Effect - Below Median		-0.026	-0.018	-0.026
Std. Effect - Above Median		-0.003	-0.010	-0.010

Notes: The dependent variable is the gross growth rate of capital between 2006 and 2010. Medians in columns 2, 3, and 4 are calculated on the distribution of the residuals of the specified variable, after regressing it on fixed assets, squared fixed assets, ROA, sales over assets, and the investment rate, all measured in 2006. Controls include fixed assets, squared fixed assets, ROA, cash-holdings over assets (except in column 2), sales over assets, and the investment rate, all measured in 2006. Bootstrapped standard errors based on 200 replications in parentheses with *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table 5: OLS and IV Estimates of the Effect of a Shock to the Growth Rate of Credit Granted on the 2007-2010 Investment Rate

	(1) OLS	(2) IV	(3) IV	(4) IV
Credit Growth	0.506***	1.426**		
	(0.038)	(0.559)		
Jimenez et al. (2011) Instrument			1.602**	
			(0.784)	
Long-Term Credit Growth				0.781***
				(0.276)
Controls	Y	Y	Y	Y
Sector FE	Y	Y	Y	Y
Province FE	Y	Y	Y	Y
No. of Obs	28218	28218	28218	25398
F-test on Excl. Instrument		43.966		55.889
Std. Effect	0.133	0.376	0.018	0.437

Notes: The dependent variable is the gross growth rate of capital between 2006 and 2010. Controls include fixed assets, squared fixed assets, ROA, cash-holdings over assets, sales over assets, and the investment rate, all measured in 2006. Heteroskedasticity robust standard errors clustered at the main bank and sector levels in parentheses with *** p<0.01, ** p<0.05, * p<0.1.

Table 6: Robustness Checks

	(1)	(2)	(3)	(4)
Panel A				
	Trimming at 5%	Winsorizing at 5%	Winsorizing at 10%	Add Discontinuing Firms
Exposure	-5.164***	-6.722**	-3.062***	-0.271**
	(1.637)	(3.182)	(1.118)	(0.120)
Controls	Y	Y	Y	Y
Sector FE	Y	Y	Y	Y
Province FE	Y	Y	Y	Y
No. of Obs	30650	32128	32128	29461
Exposure - St. Effect	-0.024	-0.016	-0.024	-0.005
Panel B				
	Sector-Province FE	Weighted by Sales	Main Bank FE	Net Interbank Exposition
Exposure	-1.043*	-1.683***	-1.132**	-1.254***
	(0.549)	(0.489)	(0.562)	(0.278)
Controls	Y	Y	Y	Y
Sector FE	N	Y	Y	Y
Province FE	N	Y	Y	Y
Sector-Province FE	Y	N	N	N
No. of Obs	29112	29110	29112	28780
Std. Effect	-0.018	-0.045	-0.019	-0.017

Notes: The dependent variable is the gross growth rate of capital between 2006 and 2010. Controls include fixed assets, squared fixed assets, ROA, cash-holdings over assets, sales over assets, and the investment rate, all measured in 2006. Heteroskedasticity robust standard errors clustered at the main bank and sector levels in parentheses with *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table 7: Estimate of the Effect of the Credit Crunch on the 2007-2010 Investment Rate Controlling for Investment Expectations

	(1)	(2)	(3)	(4)
	Expected Growth Rate of Investments (%)	Growth Rate of Gross Capital		
	OLS	OLS	OLS	IV
Exposure	4.805	-0.816*	-0.903**	
	(3.170)	(0.460)	(0.436)	(0.428)
Exp. Growth of Inv.			0.011**	0.007**
			(0.004)	
Credit Growth				0.762*
				(0.003)
Controls	Y	Y	Y	Y
Sector FE	Y	Y	Y	Y
Province FE	Y	Y	Y	Y
No. of Obs	996	996	996	996
F-test on Excl. Instrument				26.45
Std. Effect		-0.030	-0.032	0.491

Notes: The dependent variable in (1) is the expected growth rate of investment between 2006 and 2007; the dependent variable in (2) to (4) is the growth rate of gross capital between 2006 and 2010. Controls include fixed assets, squared fixed assets, ROA, cash-holdings over assets, sales over assets, and the investment rate, all measured in 2006. Heteroskedasticity robust standard errors clustered at the main bank and sector levels in parentheses with *** p<0.01, ** p<0.05, * p<0.1.

Table 8: IV estimates of the Effect of a Shock to the Growth Rate of Credit Granted on Additional Variables - Year 2010

	(1) Value Added	(2) Employment	(3) Labor Cost	(4) Intermediate Expenditures	(5) Trade Credits	(6) Trade Debits	(7) Growth of Trade Credits over Sales
Credit Growth	0.284***	0.135***	0.156***	0.112*	0.496***	0.182	0.214**
	(0.073)	(0.051)	(0.039)	(0.064)	(0.062)	(0.159)	(0.091)
Controls	Y	Y	Y	Y	Y	Y	Y
Sector FE	Y	Y	Y	Y	Y	Y	Y
Province FE	Y	Y	Y	Y	Y	Y	Y
No. of Obs.	28066	17486	28976	27637	27522	25359	
F-test on Excl. Instrument	120.979	79.338	99.813	88.913	142.388	86.620	145.041

Notes: Controls include fixed assets, squared fixed assets, ROA, cash-holdings over assets, sales over assets, and the investment rate, all measured in 2006. Heteroskedasticity robust standard errors clustered at the main bank and sector levels in parentheses with *** $p<0.01$, ** $p<0.05$, * $p<0.1$.