Banks, Sovereign Debt and Capital Requirements

Author: Filippo De Marco

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Filippo De Marco

Advised by Professor Fabio Schiantarelli

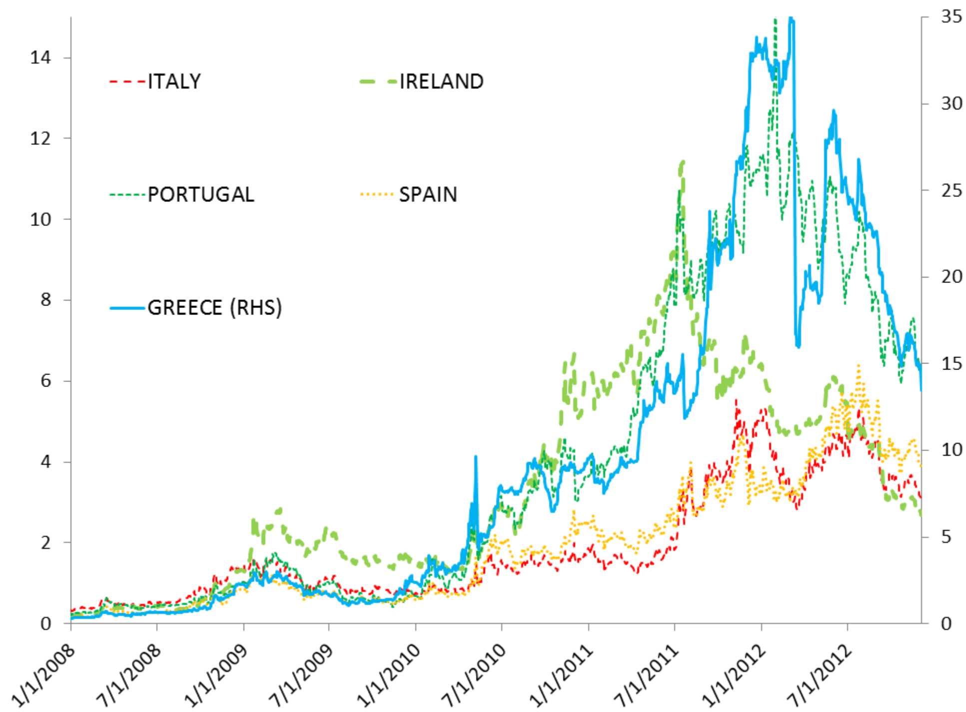
Abstract

In the aftermath of the Great Recession of 2007–2009, Europe has been grappling with both a debt and a banking crisis, which caused a prolonged recession and on–going stagnation in some countries of the Eurozone. The distinctive feature of the European crisis, compared to the global recession that originated in the United States, is that it emerged as sovereign debt crisis and later evolved into a banking crisis, finally affecting the real economy. The banking and sovereign crises are heavily intertwined because of the interplay between banks and sovereigns in Europe. In fact, the so–called *bank–sovereign nexus* works both ways: not only banks hold large amounts of sovereign debt, especially from the domestic government, but also European governments retain a significant presence in the domestic banks' ownership. The adverse feedback loop is reinforced during a sovereign debt crisis, as banks' losses from sovereign debt further exacerbate the strain on the domestic sovereign in expectation of a future bail–out. I summarize recent developments and some of the unique features of the European crisis below.

Sovereign Debt Crisis Between 2010 and 2012, a group of European countries (Greece, Ireland, Italy, Portugal and Spain, GIIPS hereafter) experienced a large increase in the borrowing costs on their respective public debt. Figure 1 below plots the evolution of the GIIPS countries 10–year sovereign bond yields, as a spread over their German counterpart. From the beginning of

2010 there has been a clear deterioration of the borrowing conditions for Greece (note the different scale on the right–hand–side), which quickly spread to Ireland and Portugal, reaching a peak at the end of 2011, when also Spain and Italy were severely affected.

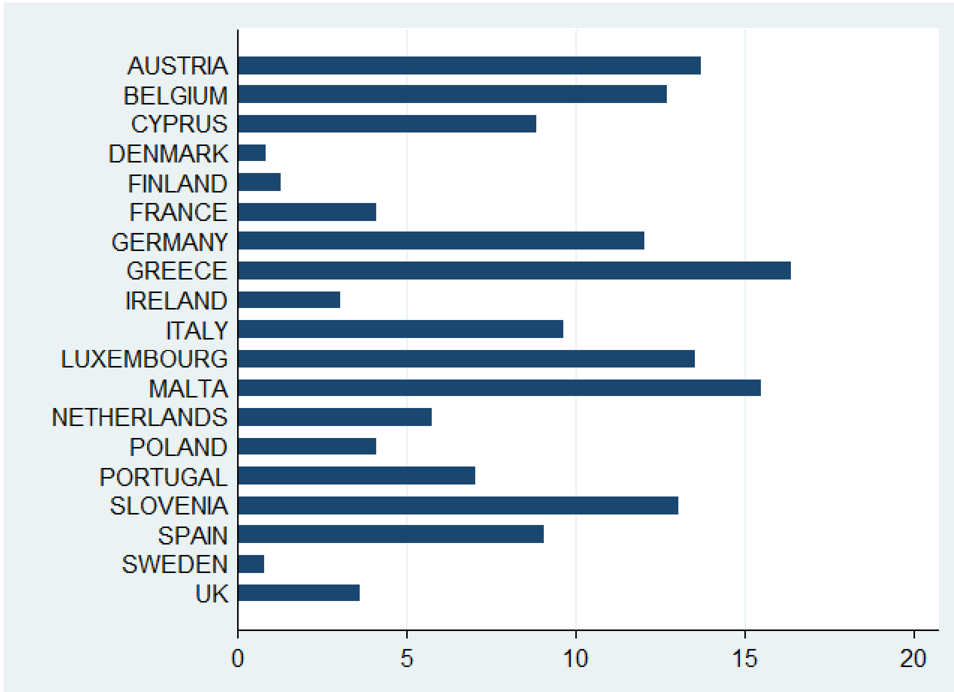
Figure 1: Sovereign Bond Yields, 10 year spread on German bonds (in %)



Note: Bond yield spreads are measured in %. Greek bond spread is on the right–hand–size (RHS) axis.

Banks' Sovereign Debt Exposure and Bank Lending European banks hold substantial amount of sovereign debt in their portfolios. Figure 2 contains the total sovereign bond exposure as a fraction of total assets for the largest European banks at the onset of European sovereign debt crisis in March 2010. On average, sovereign bond exposure represent 8% of banks' total assets in each country. This is a sizable exposure, considering that total banking assets in several countries, including the large economies, are close to or even exceed the country's Gross Domestic Product.

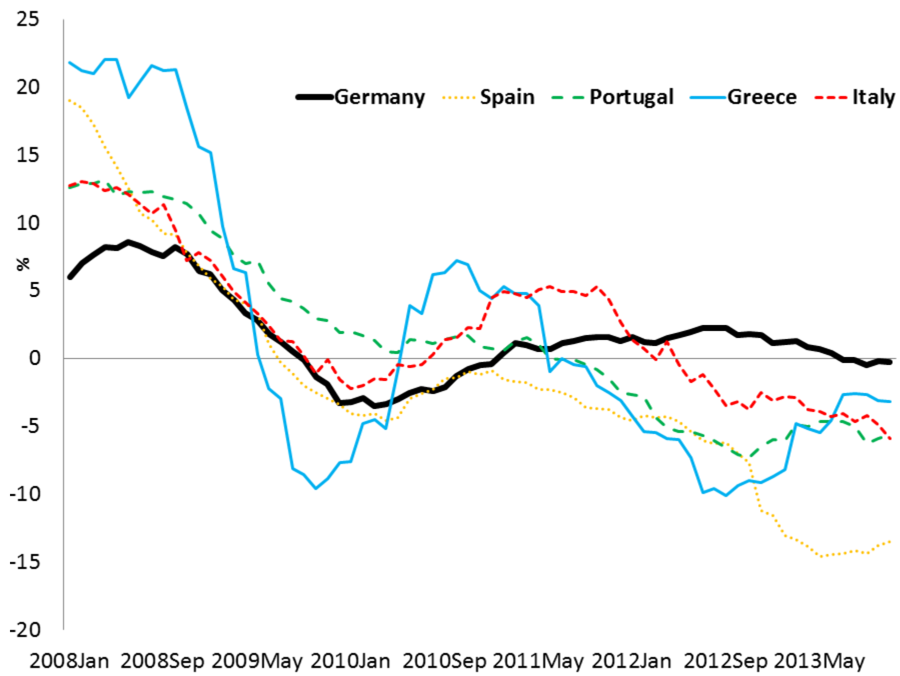
Figure 2: Total Sovereign Bond Exposure over Total Assets of European Banks, 2010Q1 (in %)



Note: Total sovereign exposure to the 30 countries of the European Economic Area (EEA 30).

At the same time, the availability of credit in the countries most affected by the increase in the sovereign debt yields rapidly declined. Figure 3 plots the growth rate of domestic loans to non–financial corporations: GIIPS countries, on average, had a much lower growth rate of lending than Germany after 2011.

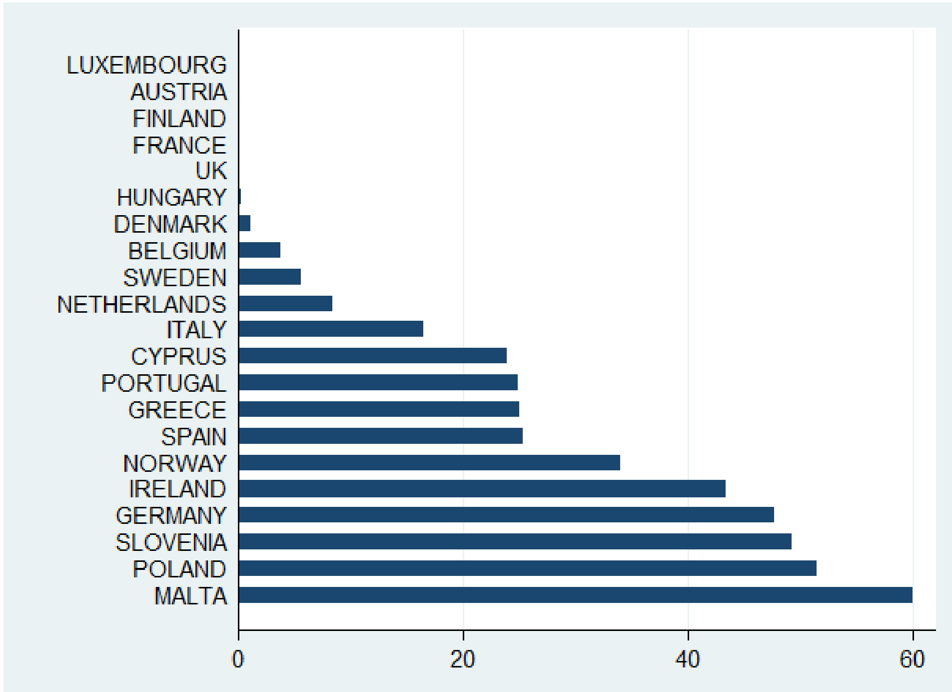
Figure 3: Growth Rate of Domestic Loans to Non–Financial Corporations (in %)



A number of commentators and policy makers have suggested that the banks' exposure to sovereign debt have contributed to the fall in aggregate lending. For example, Bundesbank governor Weidmann (2013) states that "banks that were highly exposed to strained European sovereign debt have reduced their lending to the private sector".

Government Ownership in Domestic Banks Differently from the US experience, government–owned banks are quite common in Europe. Figure 4 shows the average share of government ownership in each European country for the largest domestic institutions before the crisis (2006). Although in some countries the government was not a shareholder in any of the large banks (France, Austria and the UK), in others, including Germany, the domestic government was a large shareholder.

Figure 4: Average Domestic Government Ownership of Banks in 2006 (in %)



Capital Requirements The sovereign debt crisis has forced European regulators and policy makers to take unprecedented steps to avoid a collapse of the Eurozone. Given its fundamental role in the crisis, the European banking system has gone through an extensive regulatory scrutiny. Banks have been subject to a number of supervisory exercises, such as the European Banking Authority (EBA) Stress Tests and Recapitalization Exercises. More recently the European Central Bank (ECB), in its new role as the bank supervisor for the largest financial institutions, has done an extensive review of banks' balance sheets. A common theme across these interventions has been the need for European banks to strengthen their capital positions. The main prudential policy tool to achieve this goal are bank capital requirements, expressed as a fraction of Risk–Weighted Assets (RWA). Capital requirements are imposed by regulators to ensure that a bank has a sufficient buffer to remain solvent in case of a downturn. However, demanding higher capital requirements during a crisis may in fact exacerbate it by encouraging banks to deleverage, *i.e.* reduce RWA, to reach the target capital ratio.

The overall goal of this dissertation is to have a better understanding of the interplay between sovereign, banks and capital regulation. In my first and second chapter, I analyze the two–way feedback loop between banks and sovereigns in Europe. In particular, in the first chapter, I show that banks' sovereign debt exposures had a negative effect on credit supply during the crisis. In the second chapter I explore the role that politics may play in determining banks' exposure to sovereign debt. Finally, the third chapter investigates the effect of changing bank capital requirements for the firms that borrow from the affected banks.

Has the sovereign debt crisis induced a credit crunch, *i.e.* a reduction in credit supply, in Europe? In my first chapter, a solo–authored work, I argue that indeed this has been the case. Using detailed bank–level exposure from the EBA Stress Tests, I am able to calculate the *mark–to–market* losses that banks had on their sovereign bond portfolios. These bank–specific losses are then used to explain both the fall in lending at the bank level and the interest rate charged on syndicated loans, which are loans to large, non–financial corporations. In particular, I find that for a 1% loss of sovereign debt–to–total assets ratio, the growth rate of credit is 4% lower and, on average, interest rate charged are 2% higher. Importantly, these results are not driven by aggregate credit demand at the country level, which is also likely to fall after the sovereign shock, or by sector specific credit demand. The data suggest that the reason why these losses matter for credit is through an effect on banks' cost of funding, rather than through a negative impact on bank regulatory capital.

Does politics play a role in determining banks' exposures to domestic government debt? In the second chapter, a joint work with Marco Macchiavelli, we find that a bank where the domestic government is a relevant shareholder is associated with a disproportionate amount of domestic sovereign bonds over foreign bonds (*home bias*). We also document the extent of political pressure on the banking system during the sovereign debt crisis: political banks that were recapitalized by their respective governments increased the exposure to domestic sovereign bonds relative to foreign ones. This effect is mostly present for banks belonging to the GIIPS than for political banks located in other European countries. These results shed light on the importance of understanding the fundamental causes of the bank–sovereign nexus.

Finally, in my third chapter, a joint work with Tomasz Wieladek, we study the effect of changes to UK bank-specific capital requirements on Small and Medium Enterprises (SME) asset growth between 1998–2006. We focus on the UK experience in the pre–crisis period because of its unique regulatory regime on bank capital requirements. In fact, UK regulators imposed bank–specific and time–varying capital requirements to address risks that were not accounted for in Basel I, the standard international regulatory framework at the time. This unique policy experiment is ideal to test the impact of bank capital requirements while still being informative on the potential effects of regulatory intervention during a crisis, when presumably the effects would be even larger. We find that there are negative effects of changing capital requirements on SME growth rate of assets at the beginning of a *new bank–firm relationship*, but this effect declines over time. Shocks to the monetary policy rate also have an independent negative effect on SME asset growth and can reinforce the negative impact of increasing capital requirements. These results are important to understand the transmission channels at the firm level of prudential policy tools, an area yet relatively unexplored in the literature.

Contents

Chapter 1

Bank Lending and the Sovereign Debt Crisis

Introduction

This paper examines the effect of the sovereign debt crisis on the European credit crunch through its effect on banks' balance sheets. Banks' holdings of risky sovereign debt are substantial: Figure 1.1 shows that, at the beginning of 2010, the 90 European banks that participated in the European Banking Authority (EBA) Stress Test had a total exposure of €750 bn. to the sovereign debt of the periphery, the so–called GIIPS countries (Greece, Ireland, Italy, Portugal and Spain). These exposures are, on average, 7% of total banking assets of GIIPS banks [1] and 2–3% of the assets of banks in the core countries (France and Germany) [2], as Figure 1.2 shows. In 2010–2012, rising

[1] Most of GIIPS debt held by GIIPS banks is *domestic*. European banks, especially in the GIIPS but also in Germany, have a large *home bias*, as documented, among others, in Acharya and Steffen (2013), Battistini, Pagano and Simonelli (2013) and De Marco and Macchiavelli (2014).

[2] It should be noted, however, that French and German banks are very large compared to the size of their own economy. The top 5 French banks, for example, have total assets that are more than twice the size of the French economy. GIIPS exposures of French and German represent roughly 5% of their respective countries' GDP.

yields in the sovereign bond markets for these countries caused a devaluation of the stock of sovereign bonds held in banks' balance sheets. At the same time, the growth rate of bank credit to European customers rapidly declined. Figure 1.3 plots the evolution of bank loans to domestic non–financial firms aggregated by country. The slowdown following the 2007–2009 global recession is common to all countries, but the heterogeneity in growth rates of credit since 2010 is clear. Most GIIPS countries have lower growth rates than in the core, especially in the final part of the sample [3]. Moreover, the average interest rate spreads for corporate borrowers in the periphery has been significantly higher than in the core since 2009–2010 (Figure 1.4). The fact that loan volumes decreased while loan prices (interest rates) increased suggest that, at least at an aggregate level, the sovereign debt crisis had a negative effect on the supply of credit. In this work, I will show that the same argument holds at the micro level: losses in banks' sovereign portfolios have induced a reduction in the supply of credit from banks, both on the quantity and the price of credit. The identification will come from the cross–sectional differences in banks' sovereign exposures.

My results suggest that European banks that were more exposed to the sovereign shock significantly reduced credit supply during the crisis. In particular, a 1% increase in losses-over-total assets ratio decreases the growth rate of domestic loans around 4%. The result is not driven by country specific credit demand, even accounting for different categories of borrowers between large and small banks, or by the own domestic sovereign exposure for GIIPS banks that experienced very large shocks to their balance sheets. In this respect, I also acknowledge that sovereign exposures are endogenous and I can instrument the domestic sovereign exposure with the level of *political ownership* in each bank. In De Marco and Macchiavelli (2014) we find that banks' home bias in sovereign bonds can in part be explained by the amount of shares owned by the respective government in each bank *pre–crisis*, in 2006. We believe that banks' political ownership is fairly exogenous, as it

[3] Unexpectedly, this is not true for Italy, at least until the end of 2011. German credit growth actually fell behind Italy's in 2010–2011, underlying the difficulties of the banking sectors in the core too.

is determined by historical and political factors. If politics has a direct influence on bank lending, the IV would not be valid because exclusion restrictions are violated. I believe that this could actually work against me as political banks are generally encouraged to lend more, not less, during a crisis (Coleman and Feler (2013)). Also, political banks could have more incompetent managers or be lending to specific set of borrowers that fare worse than others during the crisis. However, if that were the case, government–owned banks should have a higher share of non–perfoming loans and lower profitability during the crisis, and they do not (see Table 1.20 in the Appendix). Also, from EBA data for the end of 2010, it does not seem that government–owned banks have a significantly different credit portfolio compared to private banks. The only difference is that non–GIIPS government–owned banks tend to lend more to domestic public sector entities (see Table 1.21 and 1.22 in the Appendix).

Moreover, to provide further evidence for a credit supply shock, I show that the European cross–border groups cut foreign loans, i.e. loans issued by their international subsidiaries, by the same amount as domestic loans. I find evidence that credit supply tightening also occurred through loan prices: loan interest rate spreads on syndicated loans are 40 to 65 basis points (bps.) higher for a one standard deviation increase in the sovereign shock (20 bps. shock), even after controlling for sector specific credit demand and borrower characteristics.

Why do sovereign losses matter for credit? There are two main channels: the *capital channel* and the *funding channel*. According to the first, banks with large sovereign losses may fall below the minimum regulatory capital ratio. Given that equity is a relatively costly source of finance and banks may be reluctant to issue new shares at a time of low stock prices, banks would find optimal to cut off loans. The funding channel indicates instead that banks with sovereign losses face a higher cost of funding, either because they pay higher rates on interbank repos (Boissel et al. (2014)), where government bonds are the preferred source of collateral or simply because a bank with higher losses is perceived as more risky by external investors. I find evidence that the second channel is more important than the first. In fact, I show that banks relying heavily on short term funding cut lending by more the higher the sovereign losses, whereas banks with low levels of regulatory capital, defined as Tier1 Equity over Risk–Weighted Assets (RWA), were not more significantly affected than other banks. In particular, this seems to be true for *regulatory* capital, expressed as a fraction of RWA, rather than for *effective* capital, expressed as a fraction of Total Assets rather than RWA: banks with low leverage ratios engaged in asset shedding more than other banks. This may indicate that sovereign losses indeed induced banks to de–leverage to ease investors' concerns about the solvency of the institution, and not as a consequence of regulatory action. I interpret the absence of a regulatory capital channel in terms of *forbearance* by the EBA in enforcing capital requirements in a time of crisis: banks were given plenty of leeway into how they could address the capital shortfall (capital requirements were effectively not binding). Also, since banks do not in fact mark–to–market their sovereign bond exposures [4] as I do in my empirical exercise, it is not surprising that the effect of sovereign losses on lending does not come through an impact on regulatory capital. The funding channel is instead the result of *market discipline*, where banks with below average collateral are either shut out of the market completely or face higher rates for wholesale funding. As additional evidence for the funding channel, I study bank–specific cost of funding using data on bond issuance from SDC Platinum. I find that coupon rates on fixed–rate notes issued by European banks increase with banks' sovereign losses, even after controlling for country–time factors and security specific characteristics.

The dataset is the result of a merger of different data sources. I construct the bank specific

[4] On average, 40% of sovereign exposures are in the Hold–to–Maturity (HTM) and 40% in the Available–For–Sale (AFS) books and the rest in the Fair–Value–Option (FVO) and the Held–For–Trading (HFT) books. Exposures in the HTM are valued at historical prices. Exposures in the AFS are marked–to–market, but banks can then net out valuation losses or gains in the calculation of regulatory equity (the so–called Prudential Filters). The application of Prudential Filters varies by country, but it implies that in many cases even AFS exposures do not impact regulatory equity.

shock, *i.e.* potential losses on sovereign debt holdings, using data on exposure to sovereign debt for the European banks participating in the European Banking Authority (EBA) Stress Tests and Recapitalization Exercises. The granularity of the EBA data, containing individual bank–level exposure to each European country and its maturity structure, allows me to exactly calculate the capital loss (profit) caused by the deterioration (appreciation) of existing bond holdings on each bank's balance sheet during the sovereign crisis. I then match the EBA exposure data with balance sheet information and explore the significance of the bank–specific sovereign shock in explaining loan growth, controlling for aggregate credit demand conditions at the country level with a set of country–time fixed effects. The key identifying assumption is that there are no systematic differences in unobserved borrowers' characteristics between the most exposed banks and the least exposed ones once time–varying, country–level unobservables are accounted for [5]. Note that I can somewhat relax this assumption in the analysis of loan interest rate spreads. Syndicated loans from LPC DealScan, in fact, contain the identity, industry and location of corporations borrowing in the syndicated loan market. This allows me to control for credit demand by introducing country–industry–quarter fixed–effects, so that I am comparing loans made to borrowers in the same country, industry and quarter by banks with different sovereign exposures. Moreover, for those borrowers that can be linked to Compustat, I introduce firm–level controls to further control for credit demand. Finally, bank bond issuance is from SDC Platinum, a rich database on equity, bond, syndicated loans issuance and M&A for both financial and non–financial firms.

This paper contributes to a recent, but active area of research, both theoretical and empirical, that studies the relationship between sovereign and banking crises. Whereas most applied papers look at cross–financial linkages between the two (Acharya and Steffen (2013) on bank stock returns, De Bruyckere et al. (2012) on sovereign contagion, P. Kallestrup et al. (2012) on bank and sovereign

[5] This point is made more rigorously in a simple model of bank lending in Section 3.

CDS, Colla and Bedendo (2014) and Augustin et al. (2014) on corporate CDS) or at the relationship between sovereign risk and bank bailouts (Acharya, Drechsler and Schanbl (2014), Greenwood et al. (2012)), few have focused on the real effects of the sovereign debt crisis. Gennaioli et al. (2013, 2014) analyze the effects of sovereign *defaults* on bank lending, thus this paper is different in that, other than Greece, there has been no sovereign default in Europe during the crisis. To my knowledge, only Bofondi et al. (2013) and Popov and Van Horen (2014) look at the effects of the sovereign debt crisis on credit supply in Europe as I do (Acharya, Eisert, Eufinger and Hirsch (2014) analyze the real effects on borrowers using syndicated loan data while Correa, Sapriza and Zlate (2013) look at lending by US branches of European banks). The main contributions of this paper compared to the latter are the following. First, I investigate the effect on both credit quantity and prices. As far as I am aware, this is the first paper that examines both these outcomes for European banks during the sovereign debt crisis. Second, I attempt to shed light on the mechanisms as to why sovereign losses matter for bank credit. I claim that risky sovereign bonds exposure matter because they increase bank cost of funding, rather than by decreasing regulatory capital. Third, I specifically address the endogeneity of sovereign exposures by instrumenting the domestic exposure with the degree of political influence, as measured by government ownership, in the domestic banking system. This is a novel approach in the literature.

The rest of the paper is organized as follows. Section 1 contains a brief review on the empirical literature on credit supply shocks and the sovereign debt crisis. Section 2 describes the data and the construction of the bank specific sovereign loss, while Section 3 outlines the empirical methodology. Section 4 presents the regression results and Section 5 discusses the possible channels at work. Finally Section 6 concludes.

1.1 Literature Review

The academic literature on credit supply shocks faces significant empirical challenges for the identification of a causal effect. In fact, data on loans from banks' balance sheets are the outcome of credit supply and demand and thus it is hard to disentangle between the sources of the shock. Traditionally, the literature has either exploited clever identification schemes (Kashyap et al. (1993), Kashyap and Stein (2000), Ivashina and Scharfestein (2010)) or examined specific institutional/quasi–experimental settings that allowed a clear separation from the two (Jayaratne and Strahan (1996) on the finance–growth nexus; Paravisini (2008) on financing frictions; Peek and Rosengren (1997, 2000) on the international transmission of bank shocks; Rocholl et al. (2011) on retail credit in Germany during the financial crisis) or, finally, having access to detailed firm–bank relationship data, included firm–period fixed effects to fully control for credit demand (Khwaja and Mian (2008) seminal work; Jimenez et al. (2012, 2014) for Spain; Bofondi et al. (2013) for Italy; Iyer et al. (2014) for Portugal).

Bofondi et al. (2013) provide evidence that foreign banks operating in Italy, mainly German and French groups, tightened credit volumes less and charged lower interest rate than Italian banks during the sovereign debt crisis. However, they do not find any differential impact for banks that are more exposed to GIIPS debt. I emphasize instead that banks with higher sovereign exposure and thus higher losses, were both cutting lending volumes by more and charging higher interest rates.

Popov and Van Horen (2014), on the other hand, is more closely related to my work. They also use the EBA stress test data to show that non-GIIPS European banks with higher exposure to GIIPS debt decreased the volume of syndicated loans at the country–borrower level by more than less exposed banks. In a robustness check, I also confirm their result on syndicated loans volume using my measure of the sovereign shock. Compared to their work, I analyze data from banks' balance sheets, that includes loans to small and medium enterprises and households (syndicated loans represent only 10% of aggregate lending in Europe, according to ECB MFI statistics). I also use syndicated loan data to investigate a "price channel" of the sovereign debt crisis through which higher sovereign losses translate into higher syndicated loan interest rates. Moreover, I provide evidence for the mechanism at work, showing that sovereign losses matter for banks' cost of funding rather than for regulatory equity.

This paper is also related to the literature on the role of politics in the banking system. The predominant, political economy view considers state owned enterprises to be a source of inefficiency (Shleifer and Vishny (1994)), maximizing the individual goals of politicians, such as over–employment and financing favored enterprises, rather than their own value. This view has the support of some influential empirical evidence: La Porta et al. (2002) show that countries with a higher share of government owned banks have lower growth rate of GDP per capita; Sapienza (2004) shows that state–owned banks in Italy charge lower interest rates than privately owned banks and Dinç (2005) shows that government–owned banks increase lending in election years, suggesting inefficient allocations. Yet, other papers offer a different view: Andrianova et al. (2008) claim that the results in La Porta et al. (2002) are not robust to appropriately controlling for the quality of institutions [6]. Coleman and Feler (2014) suggest that government–owned banks in Brazil are actually better equipped to lend more, rather than less, during a crisis. Also, they show that the quality of lending by state–owned banks actually improved during the crisis. The strongest evidence in favor of the political economy view in Sapienza (2004), *i.e.* the fact that the stronger the political party affiliated to the bank, the lower the interest rate charged, is actually economically small. In Dinç (2005), the effects are present only for emerging countries, with dubious quality of institutions. Also, he mentions that although government owned banks increase their lending in election years, the share of loans in total assets is not higher than at a privately owned banks, but government securities as a fraction of total assets are about 50% higher. Thus the main difference between

[6] Actually, in La Porta et al. (2002) paper itself, the significance of the coefficient is largely dependent on the set of controls.

government and privately owned banks seem to be the exposure to sovereign debt, as in this paper.

1.2 Data

The final dataset is the result of the merger of different data sources. The master dataset consists of the EBA sovereign exposure data collected during the "EU–wide Stress Test" and "Recapitalization Exercises". Specifically, the EBA, in an effort to enhance transparency and restore confidence in the financial system, decided to disclose on its website bank–by–bank result for both the 2010 and 2011 Stress Test Results[7] and the so–called 2011 and 2012 Recapitalization Exercises. These exercises contain information on the capital composition (including government's support measures), credit risk exposure and, most importantly, sovereign debt exposure to each of the 30 members of the European Economic Area (EEA 30) at different maturities for all the participating banks. The 2010 and 2011 Stress Tests sample consists of 90 European banks, covering more than 60% of banking assets in Europe and at least 50% in each Member State. In the 2011-2012 Recapitalization Exercises, the sample is restricted to around 60 banks, because smaller, non–cross border institutions were excluded [8]. In conclusion, the EBA exposure data offer an unbalanced, bank–level panel of sovereign exposure data, at irregular frequency (2010Q1, 2010Q4, 2011Q3, 2011Q4 and 2012Q2). For all the dates except in 2010Q1, a detailed breakdown by residual maturity, from 3 months to 15 years, is also provided [9].

A possible concern with this type of supervisory data is that banks may have been "window dressing" their balance sheets around the stress tests' reporting dates. There is some evidence (Acharya

[7] Data for the 2010 Stress Test, with data as of March 2010, were published on the former banking authority website, CEBS. A link to the sovereign exposure for this date can be found at the Peterson Institute for International Economics `http://www.piie.com/blogs/realtime/?p=1711`

[8] Most of the excluded banks are from Greece or Spain. The six Greek banks, present in the Stress Tests, were under restructuring and IMF monitoring by the time of Recapitalization Exercises. Most of the regional savings bank in Spain were excluded too.

[9] I will extrapolate the maturity structure for 2010Q1 exposure from the 2010Q4 data, by using the same proportions.

and Steffen (2013), Acharya, Engle and Pierret (2014)) that some banks effectively "gamed" the second stress test by reducing their GIIPS exposure right before the reporting date and increasing it following the disclosure of the results. However, even if this behavior has indeed been common practice among banks, it means that the EBA data understate the true amount of sovereign debt on banks' balance sheets, so that any negative effect of the sovereign shock on bank loans would be underestimated.

Banks' exposures are then matched with balance sheet information either from Bankscope or hand–collected (and cross–checked) from banks' annual reports. Since balance sheet data have a yearly frequency, I end up eliminating the mid–year EBA exposures on 2011Q3 and 2012Q2. On the other hand, I impute the 2010Q1 sovereign exposure to the beginning of the year and I match it with the end–of–year 2009 balance sheet data. The EBA sovereign exposure data are provided at the group level for the reporting banks; thus, the matching with the balance sheet data has to be done carefully. Balance sheet controls in the regressions are kept at the same "highest" level of consolidation, but the outcome variables, either domestic or foreign loans, need to be adjusted depending on the size of the bank. For smaller, local banks, domestic loans come from the unconsolidated statements, whenever possible, or the consolidated group statement otherwise. For the larger, global bank groups, domestic loans either consist of the unconsolidated figures of the parent bank or, when no unconsolidated statements were available, they are computed as the consolidated total minus the loans of the foreign subsidiaries belonging to the group. For the cross–border groups only, I am able to construct a series for foreign loans using the unconsolidated loans of the international subsidiaries. In this case, I end up with 36 cross–border banks out of the original 90 banks of the EBA sample, yielding a total of 140 subsidiaries (see Appendix for a list of banks).

Table 1.1 contains summary statistics for the EBA sample matched with balance sheet characteristics at the group level. Banks in the sample are quite large, with average (median) Total Assets of €390 bn. (€135 bn.) with banks above the 90^{th} pct. exceeding €1 tn. On average these assets are funded through customer deposits for 40% and 10% through other, short–term funding sources. They also seem to be well capitalized according to regulatory capital, as even banks in the 10^{th} pct. have a capital ratio of 8% (Basel I benchmark) and more than half are well above 10%. In terms of the effective capital ratio, expressed as a fraction of total assets, the capitalization number are quite different. The numbers are halved on average, with banks in the 10^{th} pct. having an implied leverage of 50 or more. The difference in the distribution of the two capitalization measures underlines the flexibility of risk–weights in computing regulatory capital. Profits and the growht rate of domestic lending are on average negative and non–performing loans (NPL) are quite high, as expected in a crisis period. Finally, 50% of banks' balance sheet on average consist of loans, while sovereign bonds take up another 9%, with exposure to GIIPS taking 3.7% on average.

I also merge the EBA sample with LPC DealScan that contains loan–level information on interest rate spreads. The largest EBA banks are especially active in the syndicated loan market, with at least one of the top 10 European banks being present in more than 75% of the syndicated loans granted by the EBA sample banks over 2009-2012. Interest rate spreads are available at the tranche (facility) level, so that the relevant panel–id variable is the bank–facility pair at a quarterly frequency. Moreover, I could match some of the borrowing firms in DealScan to accounting data

Table 1.1: Summary Statistics EBA Sample

Variable	Obs	Mean	Std.Dev.	10^{th}	25^{th}	50^{th}	75^{th}	90^{th}
TotAss (€bn.)	232	388.08	576	20.56	45.7	135.5	405.3	1,160
Deposits/TA	230	0.432	0.16	0.233	0.307	0.439	0.53	0.63
STFund/TA	234	0.109	0.089	0.018	0.046	0.09	0.145	0.203
Tier1/RWA	228	0.116	0.039	0.083	0.097	0.112	0.131	0.152
Tier1/TA	232	0.05	0.028	0.02	0.034	0.047	0.064	0.078
Profits/TA in (%)	232	-0.097	1.978	-1.16	0.02	0.3	0.55	0.829
NPL/Loans (in %)	225	6.95	5.51	2.05	3.27	5.66	8.93	13.64
ΔLoans domestic	233	-0.011	0.1	-0.1	-0.042	-0.007	0.029	0.074
Loan/TA	230	0.505	0.21	0.236	0.333	0.51	0.681	0.77
TotSov/TA	201	0.091	0.058	0.027	0.047	0.083	0.122	0.153
GIIPSExp/TA	201	0.037	0.048	2.40E-07	0.003	0.012	0.065	0.097

available from Compustat using the link file provided by Chava and Roberts (2008). I exclude corporate borrowers in the financial, insurance and real estate sectors [10] Table 1.2 below provides some summary statistics for these two dataset mergers. Syndicated loans are large, with a mean (median) of $430 mil. ($150 mil.), have an average maturity of 5 years, average all-in drawn spread over the reference rate (Libor or Euribor) of 300 basis–points and attract an average (median) of 4.2 (3) participant banks, including 3 (2) Lead Arrangers. The DealScan-Compustat sample is broadly consistent with the DealScan-EBA sample and, although loans are on average twice as large, the spreads, maturity and number of banks are similar. Borrowing firms are also very large, with average (median) assets of $25.9 bn. ($8 bn.), average profit margin (EBITDA/Sales) of 0.22 and average leverage ratio of 0.36.

Table 1.2: DealScan Summary Statistics

	DealScan-EBA Sample					
	N	Mean	Std.dev.	10^{th}	50^{th}	90^{th}
Loan characteristics						
Loan Amount ($ mil.)	11,795	430	960	18	150	1,140
All–in drawn spread (bps.)	7,041	301.2	162.4	115	275	500
Maturity (months)	11,331	65.14	50.32	18	60	102
Number of participants	11,810	4.2	4.3	1	3	9
Number of arrangers	11,810	3	3	1	2	7
	DealScan-Compustat Sample					
	N	Mean	Std.dev.	10^{th}	50^{th}	90^{th}
Loan characteristics						
Loan Amount ($ mil.)	1,386	984	1,440	100	513	2,120
All–in drawn spread (bps.)	1,208	248.8	133.7	100	225	425
Maturity (months)	1,360	55.84	23.6	30	60	78
Number of participants	1,386	4	4.5	1	2	13
Number of arrangers	1,386	2.6	3	1	1	6
Borrower characteristics						
Tot.Ass.($ bn.)	833	25.9	51.9	0.96	8	56
EBITDA/Sales	792	0.22	0.72	0.05	0.2	0.49
Leverage	805	0.36	0.24	0.13	0.35	0.63
Investment/Assets	667	0.015	0.017	0.003	0.01	0.029

Finally, bond yields are taken from Bloomberg. For the construction of the sovereign shock, I keep only maturities longer than or equal to 2 years because shorter maturities, 3–months (3M)

[10]SIC codes between 6000 and 6999

and 1–year (1Y), contain a lot of noise, have missing values and do not matter as much for the computation of sovereign losses (short duration). Also, due to data availability, I can match bond yields to the sovereign exposures of 17 countries only out of the original EEA 30: Austria (AT), Belgium (BE), Germany (DE), Denmark (DK), Spain (ES), Finland (FI), France (FR), Greece (GR), Hungary (HU), Ireland (IE), Italy (IT), Netherlands (NL), Norway (NO), Poland (PO), Portugal (PT), Sweden (SE) and the United Kingdom (UK).

In conclusion, the final dataset contains a bank–level panel with around 90 European banks in 20 countries, matched with both balance sheet variables over 4 years (2009-2012), individual bank–by–bank sovereign exposure to 17 countries, at five different maturities (2Y, 3Y, 5Y, 10Y, 15Y) and syndicated interest rate spreads made by 74 EBA banks over the 2010–2012 period.

1.2.1 The Bank–specific Sovereign Shock

The advantage of the EBA data is that it provides a detailed picture of bank exposure to central governments of all 30 countries of the EEA with a breakdown by maturity, from 3M to 15Y. The coupon rate for the bonds in the sovereign portfolio is not known. Thus, in order to calculate the losses I have to make some assumptions on the duration of these bonds. The maturity detail is also important for the calculation of losses, as bonds with longer maturities (duration) will have a larger impact.

I construct a bank–specific sovereign shock for bank b at time t, $SovShock_{b,t}$ as:

$$SovShock_{b,t} = \sum_{s=1}^{S} \sum_{m=2Y}^{15Y} Duration_{s,m,t} * \Delta yield_{s,m,t} * \frac{Exposure_{b,s,m,t-1}}{Total\ Assets_{b,t-1}} \tag{1.1}$$

where s is the sovereign country whom bank b is exposed to; m is the residual debt maturity, in years, and t is the end of year t, from 2010 to 2012. Essentially, this shock represents the *potential* capital loss (gain) incurred by bank b during year t because of the depreciation (appreciation) of sovereign bonds. In other words, it is the marked–to–market value of the total exposure to sovereign bonds on banks' balance sheets. Although banks do not necessarily need to mark–to–market these exposures, especially if they are held in the Hold–to–Maturity (HTM) banking book, this measure is meant to capture expected losses on sovereign bonds and identify the banks most vulnerable to the sovereign shock from the point of view of an outside investor. $SovShock_{b,t}$ is composed of several terms that I define below. Notice that, as described in the data section, I eliminate maturities shorter than 2Y and keep the exposures to 17 countries ($S = 17$). Also, since the March 2010 exposure data are disaggregated by country of exposure but not by maturity, I have assumed that the maturity structure of the sovereign portfolio has remained constant over the year and I have imputed the December 2010 maturity proportions to the March 2010 figures.

$Duration_{s,m,t}$ is the *modified duration* and it measures the percentage change in the price of a bond for a unit change in the yield–to–maturity ($yield$). If the duration is, say, 10 then the price falls by 10% for any 1% increase in the yield. Sovereign bonds are coupon bonds and to compute the exact duration one would need to know the actual coupon value. However, since this information is not available in the EBA data, I have to assume that sovereigns are either *zero–coupon* bonds or *par* bonds (where the coupon equals the yield). Since the duration is a decreasing function of the coupon, using the par bond assumption will underestimate banks' losses, whereas the zero–coupon bond will overestimate them. Therefore, my preferred measure to calculate the duration is the par bond, but the main results are not qualitatively affected by this assumption. According to the par bond assumption, the duration is:

$$Duration_{s,m,t} = \frac{1}{yield_{s,m,t}}\left(1 - \frac{1}{(1 + yield_{s,m,t})^{2m}}\right)$$

where, given the assumption of semi–annual payments, $yield_{s,m,t}$ is the semi–annual yield and thus the maturity (in years) is multiplied by 2. For the zero–coupon bond, the duration becomes:

$$Duration_{s,m,t} = \frac{2m}{1 + yield_{s,m,t}}$$

It is important to keep in mind that both measures will contain a measurement error because in reality a sovereign bond is “in between” a zero–coupon and a par bond, with the latter being a closer approximation. Thus, in either case and as long as the measurement error is white noise, the OLS estimate will be biased towards zero as in the case of classical measurement error (Greene(2012)). An average of the par and zero coupon bond should contain less measurement error: I do find that this is the case in a robustness test (Table 1.10).

The second term in (1.1), $\Delta yield_{s,m,t}$, is simply the change in the average (semi–annual) yield for the month of December (or in the last quarter) for the sovereign debt of country s at maturity m. Thus the first part of the expression, $Duration_{s,m,t} * \Delta yield_{s,m,t}$, represents the total price change over the period observed, in this case from December of year $t-1$ to year t. This is the aggregate sovereign shock which is identical for all banks. Once this is interacted with bank b exposure to each sovereign c and maturity m at the beginning of the period, $Exposure_{b,s,m,t-1}$, and it is normalized by total assets, $Total\ Assets_{b,t-1}$, the loss (or gain) for bank j on that specific bond is obtained.

Finally, by summing over each country of exposure s and each maturity m, $SovShock_{b,t}$ calculates the losses (gains) from the devaluation (revaluation) of all sovereign bonds as a percentage of total assets. Table 1.3 reports the empirical distribution of the shock in the data using both the par bond and zero–coupon assumption to compute the duration. Not only the mean and standard deviation differ substantially, but the entire distribution using the zero–coupon assumption, especially in the upper tail, is wider. This is consistent with the fact that the zero–coupon bond overestimates the losses. In both cases, banks at the bottom of the distribution actually recorded some small gains on their holdings of sovereign debt, around 0.5% of total assets[11]. In fact, some Northern European countries, to which German, Danish, Dutch and Swedish banks have considerable exposures, actually experienced a decrease in bond yields during the sovereign debt crisis (flight–to–quality).

Table 1.3: Distribution of $SovShock_{b,t}$

Percentile	Loss (if positive)/ Gain (if negative)	
	Par bond	*Zero–coupon*
10^{th}	-0.563%	-0.681%
25^{th}	-0.16%	-0.167%
50^{th}	0.018 %	0.039%
75^{th}	0.350%	0.51%
90^{th}	0.964%	1.449%
95^{th}	2.75%	5.28%
Mean	0.375%	0.848%
std.dev.	1.89%	4.247%
Obs.	240	240

Banks in the 75^{th} and 90^{th} percentile had, respectively, losses accounting for 0.35% (0.51%) and 0.96% (1.45%) of total assets using the par (zero coupon) bond assumption. These numbers are high: considering that the median capital–over–total asset ratio of around 5% over 2009-2011, losses in the top decile have the potential to wipe out almost a quarter (or third) of the book value of equity. Banks facing these heavy losses are mostly headquartered in the GIIPS countries, but in the top quartile we also also find some banks domiciled in Belgium (Dexia), Germany (Commerzbank and Hypo Real Estate) and Luxembourg (BCEE).

[11] $SovShock_{j,t}$ is positive if there are losses and negative if there are gains. This is because *duration* is defined as $-dP/dyield$

1.3 The Empirical Methodology

The baseline empirical specification is the following:

$$\Delta Loans_{b,c,t} = \beta_1 SovShock_{b,t} + \eta_b + \lambda_{c,t} + \gamma' X_{b,t-1} + \epsilon_{b,c,t} \tag{1.2}$$

estimated with either OLS (with bank fixed-effects) or Difference GMM. $\Delta Loans_{b,c,t}$ is the annual growth rate of loans granted by bank b in country c (either domestic or foreign) at the end of year t; η_b is the bank fixed-effect; $\lambda_{c,t}$ is the country–year fixed–effect that accounts for country-specific credit demand; $X_{b,t-1}$ is a vector of bank balance sheet characteristics at the beginning of the period (Tier1Ratio, Pre–Tax Profits, Customer Deposits, Non–Performing Loans and Cash, all normalized by total assets [12]). The main coefficient of interest in (1.2) is β_1: I expect $\beta_1 < 0$, so that losses from the holdings of sovereign debt, all else equal, should have a negative impact on credit growth.

Table 1.4: Mean Differences in Bank Characteristics by GIIPS exposure

	2009		2010		2011	
	diff	t-stat	diff	t–stat	diff	t–stat
TotAss (€bn.)	330.0	3.05	347.2	3.34	190.7	1.17
Dep/TA	-0.064	-1.84	-0.086	-2.745	0.044	1.129
STFund/TA	0.009	0.442	0.012	0.64	0.02	0.889
Tier1/RWA	0.012	2.4	0.014	1.8	0.005	0.584
Tier1/TA	-0.018	-3.33	-0.009	-1.59	2.40E-05	0.004
Prof/TA	-0.001	-1.43	0.002	1.46	0.011	2.386
NPL/Loans	-0.47	-0.886	-0.438	-0.56	-1.33	-0.963
Δ Loans domestic	0.008	0.298	0.022	0.729	0.062	3.74

This table presents the difference of means in bank characteristics by GIIPS exposure. In particular, for each year, it takes the difference in the average of each characteristic for banks below the median in GIIPS/Total Assets and for banks above the median.

Given that the identification comes from cross–sectional variation in sovereign exposure, it is important to describe differences in bank characteristics between more and less exposed banks, both ex–ante and during the crisis. Table 1.4 examines differences in averages between the banks

[12] Among these covariates, the Non–Performing Loans ratio controls for the average quality of the loan portfolio within each bank. This is important because we may worry that sovereign exposures towards GIIPS countries are correlated with the average loan quality held in banks' balance sheets, as exemplified by the Cypriot banks case (A&S (2013)).

less exposed to GIIPS debt (below the median) and those more exposed (above the median). It is evident that less exposed banks are much larger on average (€200-300 bn.), as the big, global banks fall in this category. Importantly, the two groups are almost identical in terms of wholesale funding, profitability and non–performing loans both before and at the beginning of the crisis (2009 and 2010). More exposed banks, as expected, becomes less profitable (1.1% lower in profits/assets ratio) and experience lower growth rate of lending (6% lower) at the peak of the crisis, in 2011. Interestingly, before the crisis, the least exposed banks have a higher regulatory capital ratio (+1.2% in Tier1/RWA), but also a lower leverage ratio (-1.8% in Tier1/TA): this may be due to the fact that large banks, that fall in the less exposed category, use internal models for the calculation of risk–weights and hence can use regulatory leverage more effectively.

Table 1.5: Mean Differences by GIIPS exposure for non–GIIPS banks

	2009		2010		2011	
	Diff	t–stat	Diff	t–stat	Diff	t–stat
TotAss (€bn.)	-83.5	-0.43	-32.16	-1.66	-150.1	-0.68
Dep/TA	0.05	0.95	0.124	2.672	0.074	1.502
STFund/TA	0.038	1.12	0.005	0.197	-0.016	-0.608
Tier1/RWA	0.024	4.64	-0.002	-0.228	-0.001	-0.119
Tier1/TA	0.018	2.6	0.026	4.151	0.018	2.745
Prof/TA	0.002	1.47	0.005	3.6	0.001	0.793
NPL/Loans	-0.05	-0.075	-0.646	-0.501	0.952	0.608
Δ Loans domestic	-0.09	-1.78	0.058	1.08	0.045	2.358

This table presents the difference of means in bank characteristics by GIIPS exposure for non–GIIPS banks. In particular, for each year, it takes the difference in the average of each characteristic for banks below the median in GIIPS/Total Assets and for banks above the median.

Table 1.5 replicates the same exercise for non–GIIPS banks only. Interestingly, non–GIIPS banks more exposed to GIIPS debt are now actually larger than the less exposed banks, however the difference is not statistically significant. Importantly, other than being less capitalized, the most exposed non–GIIPS banks are similar to the least exposed banks in terms of non–performing loans, wholesale funding and profitability before and during the crisis. Finally, non–GIIPS banks more exposed to GIIPS debt have a lower growth rate of domestic loans than the less exposed ones in

2011. Since demand in non–GIIPS countries is not correlated with the sovereign shock in GIIPS countries, this is an indication of a supply, rather than demand, effect.

When I look at the effect of the sovereign shock on *domestic* loans, the identification is particularly strong for banks headquartered in countries whose bond markets were not under pressure (non–GIIPS countries), but that nonetheless had high exposure to risky sovereign debt. In fact, the sovereign shock for these banks is plausibly exogenous with respect to domestic credit demand condition. For example, Greek sovereign problems should not affect aggregate demand conditions for German firms [13], but it would affect credit supply in Germany if its banks are highly exposed to Greek debt. On the other hand, in GIIPS countries aggregate demand conditions are probably negatively correlated with the rise in bond yields and one may worry that controlling for the country–period fixed-effects is not enough to take care of the endogeneity bias caused by the home–country exposure of GIIPS banks. I address these concerns on home–country and other endogeneity biases in several ways.

First I look at the effect of the sovereign shock on *foreign* (worldwide) loans of the largest, cross-border institutions, i.e. loans granted by the international subsidiaries of the largest banking groups. If, following a negative shock to the balance sheet of the mother bank, we observe that lending is reduced also abroad ($\beta_1 < 0$), then it must be a because of a credit supply shock.

Second, I can split the sovereign shock in two parts: one part would account for losses coming from exposure to the GIIPS, while another for losses, or gains mostly, from non–GIIPS exposure. Then, I can divide each part of the shock between GIIPS and non–GIIPS banks: if the effect of sovereign losses coming from GIIPS exposure is present also for lending by non–GIIPS banks then this is a further indication of a supply, rather than a demand shock.

Third, I can instrument the sovereign shock using the level of *government ownership* in each

[13] Germany's export to Greece are marginal, on average around 1% of total German exports over the last 10 years

bank. In De Marco and Macchiavelli (2014), we show that the large degree of home bias can in part be explained by the degree of political connections of each bank with its domestic government, as measured by the the percentage of bank shares owned by the local or national government *pre–crisis*, in 2006. I can use this instrument to predict the domestic exposure in the construction of the sovereign shock. In particular, recalling that the sovereign shock is constructed as:

$$SovShock_{b,t} = \sum_{s=1}^{S} \sum_{m=2Y}^{15Y} Duration_{s,m,t} * \Delta yield_{s,m,t} * \frac{Exposure_{b,s,m,t-1}}{Total\ Assets_{b,t-1}}$$

I instrument only the sovereign exposure part, $Exposure_{b,s,m,t-1}/Assets_{b,t-1}$, when bank b is exposed to its *domestic* sovereign [14], while I let the yield follow its actual path. The implicit assumption in doing this is that the aggregate shock ($Duration_{s,m,t} * \Delta yield_{s,m,t}$) is exogenous with respect to banks' conditions, so that it does not need to be instrumented. Hence, the IV for $SovShock_{b,t}$ is constructed as follows:

$$SovPolitical_{b,t} = Duration_{10Y,t} * \Delta yield_{10Y,t} * Political_b$$

where $Political_b$ is the share of government ownership in bank b in 2006; duration and yield are for the *home* country where bank b is located and are measured at the 10 year maturity. By construction, $SovPolitical_{b,t}$ is time–varying only insofar as the yield of a country is time–varying, however it does not contain any time–variation at the bank level. Hence, although it will not technically be absorbed by the bank fixed–effects, it is possible that there is not enough variation to estimate it precisely. An alternative to the within transformation is the first–differencing estimator, that equally removes bank fixed–effects. Thus, as additional evidence, I will also be presenting the results from

[14] Foreign exposures cannot be instrumented with domestic political ownership, hence I treat them as exogenous once the home exposure has been set. Note that the average (median) home bias across banks in sovereign bonds is 74% (80%), thus the large part of the sovereign portfolio is domestic.

a cross–sectional IV regression comparing the growth rate of lending before and during the crisis.

The instrument itself may be endogenous, *i.e.* correlated with the unobserved component in credit conditions, if we think that "political" banks are poorly managed and have low profitability. It could also be the case that political banks lend to politically connected firms who could potentially be doing worse than other firms during the crisis [15]. However, if that were the case, political banks should have a higher share of non–performing loans or lower profits than non–political banks during the crisis, but that does not seem to be the case (see Table 1.20 in the Appendix). Moreover, I can specifically look at banks' loan portfolios in December 2010. In fact, the EBA released data on the allocation of credit broken down by country of destination and by type of loan in the 2011 Stress Test only [16]. For instance, I know the amount of exposure to public sector entities, corporate, SME and commercial–real–estate by country for each bank. Tables 1.21 and 1.22 in the Appendix show that government–owned banks only have a higher share of loans to domestic public entities compared to privately owned banks. This may be a problem for exclusion restrictions for GIIPS "political" banks, because domestic public sector entities in these countries are likely to suffer more during the debt crisis. However, it is especially non–GIIPS government–owned banks that tend to lend more to domestic public entities, even though both GIIPS and non–GIIPS government–owned banks have a higher home–bias in sovereign bonds compared to private banks (Table 1.22). Thus, the main difference between non–political and political banks is the higher share of domestic government bonds for the latter group.

Finally, when I look at loan interest rate spreads (over Libor or Euribor), I can relax the assumption that there are no systematic difference in borrowers' unobservables once country–wide credit demand factors ($\lambda_{c,t}$) are taken into account. In fact, syndicated loan data from LPC Dealscan reveal the identity, location and industry of the corporate borrowers participating in this market.

[15] Although this possibility is intriguing, I found no evidence for such behavior by politically connected firms in the empirical literature.

[16] These data on Exposures At Default (EAD) are used to calculate RWA.

Therefore, I can introduce country–industry–quarter fixed–effects, so that I am comparing loans made to borrowers in the same country, industry and quarter by banks with different sovereign exposures. Specifically, in equation (1.3) I have:

$$Spread_{b,f,t} = \beta_1 SovShock_{b,t} + \eta_b + \lambda_{c,i,t} + \phi' F_{f,t} + \gamma' X_{b,t-1} + \epsilon_{b,f,t} \tag{1.3}$$

where $Spread_{b,f,t}$ is the all-in drawn spread over the Libor or Euribor of the loan extended by bank b (Lead Arranger) to firm f at quarter t [17]. $\lambda_{c,i,t}$ is a country×industry ×quarter fixed–effect where borrower f is located. The fixed–effect identification scheme is very solid here because I am comparing firms in the same industry (2 digit NAICS), in the same country at the same quarter. Moreover, for those firms that can be matched to Compustat, I can control for a set of the borrower's balance sheet variables, $F_{f,t}$. Thus in this case I am not only comparing firms within the same sector, country and quarter, but also those with similar observable characteristics [18]. Finally $SovShock_{b,t}$ is constructed at a quarterly frequency, holding the sovereign exposure fixed at the beginning of the year ad letting the (average) yield vary in each quarter. Here, I expect $\beta_1 > 0$: banks with higher losses from sovereign bonds are going to charge higher interest rates on their loans to make up for lost profitability.

Finally, note that all standard errors have been clustered at the bank–level. The key identifying assumption for consistency of cluster–robust standard errors is that there should be no inter–cluster correlation, although intra–cluster correlation is allowed (Liang and Zeger (1986)). To account for country–specific correlation among banks headquartered in the same country all models have been

[17] I use the all-in drawn spread because, according to DealScan, it also takes into account one–time and recurring fees associated to the loan, so it is a better measure of the overall cost of the loan. Since it is a spread over the benchmark interbank rates, it also nets out the effects of monetary policy. Finally, I am focusing on Lead Arrangers because I am assuming that these banks have the pricing power in each loan, but the results do not dependent on this assumption. See Section 4.2 for a detailed discussion

[18] The balance sheet variables are Leverage, Ebitda/Sales, Investment/Asset, Fixed Interest Rate Coverage, log(Assets)

run through country–time clusters, rather than bank clusters, and the results still hold. However, I have decided to present the results for bank–clustered standard errors because for consistency one needs the number of clusters to go to infinity: I have a total of around 90 bank groups, but only 40 country–time pairs in equation (1.2).

1.4 Results

1.4.1 Loan Growth

Table 1.6 reports the results for the baseline regression in (1.2) for domestic loans using the par bond assumption. The coefficient of interest, β_1, is always negative and significant at 5%. It implies that banks that were more exposed to the sovereign shock and experienced higher sovereign losses had a lower growth rate of loans. Column (1) and (2) present the results with the GMM estimator and column (3) and (4) do the same with the OLS within estimator. The choice of instruments for GMM is the following: all balance sheet variables and the sovereign shock dated $t-1$ to $t-3$. In fact, the test for first order serial correlation in the error term in the difference equation ($\Delta\epsilon_{b,c,t}$) cannot reject the null of no serial correlation: the error term is a random walk. This allows me to use variables dated at $t-1$ as instruments for the equations in difference.

Column (1) estimates the baseline model with the Difference GMM. It implies that for a 1% increase in the sovereign losses–over–asset ratio, the growth rate of loans would decrease by slightly less than 4%. In column (2) I re–estimate the model with OLS, introducing a bank fixed effect: the coefficient is quantitatively almost the same, with a multiplier effect around 4.07. The number of observations is different from one estimator to the other because the panel has $T = 3$ (2010,2011 and 2012) and thus I "lose" one cross–section in the difference GMM equations. Finally, column (3) standardizes the shock by its standard deviation, so to ease comparisons with the robustness spec-

Table 1.6: Domestic Lending and the Sovereign Shock $\Delta Loans_{b,c,t} = \beta_1 SovShock_{b,t} + \eta_b + \lambda_{c,t} + \gamma' X_{b,t-1} + \epsilon_{b,c,t}$

	GMM (1)	OLS–FE (2)	OLS–FE (3)
SovShock	-3.808^{**}	-4.067^{**}	
	(1.625)	(1.573)	
SovShock/std.dev.			-7.735^{**}
			(2.987)
Tier1/RWA	2.873^{***}	2.928^{***}	2.928^{***}
	(0.355)	(0.408)	(0.408)
PreTaxProf/Assets	5.08^{***}	5.223^{***}	5.223^{***}
	(1.354)	(1.443)	(1.443)
Impaired Loans/Assets	-2.477^{***}	-2.353^{***}	-2.353^{***}
	(0.680)	(0.695)	(0.695)
Deposits/Assets	-0.111	-0.531	-0.531
	(0.367)	(0.557)	(0.557)
Cash/Assets	-0.231	0.071	0.071
	(0.605)	(0.648)	(0.648)
N	127	217	217
N of banks	78	89	89
bank fixed effects	no	yes	yes
country×year fixed effects	yes	yes	yes
Hansen–Sargan p–val	.73		
$AR(1)$ p–val	.836		

Cluster robust *s.e.* in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note: the dependent variable is the growth rate of domestic loans. $Tier1$ is the Tier1–Capital–Ratio; $Deposits/Assets$, $Cash/Assets$, $PreTaxProfits/Assets$, $ImpairedLoans/Assets$ are, respectively: customer deposits, cash and other cash equivalents, EBT and non–performing loans all normalized by total assets. All variables are measured the beginning of the period $(t-1)$. $SovShock_{j,t}$ is the bank–specific sovereign loss normalized by total assets too. Column (1)–(2) use the Difference GMM estimator with instruments dated $t-1$ to $t-3$; column (3)–(4) use the standard OLS with bank fixed–effects. All std.err. have been clustered at the bank level.

ifications that follow [19]. In this case the interpretation of the coefficient is that for a one standard deviation shock to sovereign losses, the growth rate of loans is expected to decrease by 7.7%. Among the balance sheet variables, the relevant ones are the Tier1 Capital Ratio, the profit–to–assets ratio and impaired loans–over–assets ratio. Not surprisingly, more capitalized, more profitable banks and banks with less non–performing loans at the beginning of the year had a higher loan growth rate during the following year.

Table 1.7: Foreign Lending and the Sovereign Shock
$$\Delta Loans_{b,c,t} = \beta_1 SovShock_{b,t} + \eta_b + \lambda_{c,t} + \gamma' X_{b,t-1} + \epsilon_{b,c,t}$$

	GMM Foreign (1)	OLS–FE Foreign (2)	GMM Both (3)	OLS–FE Both (4)
SovShock	-6.061^{***}	-4.545^{***}	-4.811^{***}	-2.233^{**}
	(1.263)	(1.469)	(1.084)	(1.127)
SovShock$\times foreign_b$			0.484	-2.575^{*}
			(0.893)	(1.543)
N	242	382	369	602
N of banks	139	140	217	230
bank fixed effects	no	yes	no	yes
country×year fixed effects	yes	yes	yes	yes
Hansen–Sargan p–val	.57	.04		
$AR(1)$ p–val	.935	.897		
Bank controls: Tier1(+), Profits(+)**, NPL(-), Dep(-)***, Cash(+)				

Cluster robust *s.e.* in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note: dependent variable is the growth rate of foreign (columns (1)–(2)) or of foreign *and* domestic loans (columns (3)–(4). $foreign_b = 1$ if the bank is an international subsidiary, zero otherwise. All other bank balance sheet variables at the group level are defined as before. All std.err. have been clustered at the bank level.

Effect on Foreign Loans The results of the baseline specification explore the effect of the sovereign shock on domestic loans, *i.e.* the loans issued by the parent bank in its own country. I now also examine the international transmission of the shock through the loans issued by the international subsidiaries of the cross–border institutions present in the EBA sample. Controlling for the

[19] In some of the robustness tests I will change the construction of the sovereign shock, thus altering the whole distribution. I find the standardization with the standard deviation easy to compare across specifications

country–year fixed–effect of the country where the subsidiary is located, Table 1.7 shows that the effect of the shock is still negative and has a similar magnitude to the effect on domestic loans. In particular, column (1) implies that the GMM estimate of the effect of the sovereign shock is actually larger (*i.e.* more negative) for foreign loans than for domestic ones (-6.9% compared -4%). The OLS–FE estimate, at 4.5% is somewhat similar to the effect on domestic lending. Column (3) and (4) merge the data on foreign loans with those on domestic loans and it turns out that the difference between the two (the interaction term with a dummy $foreign_b$ equal one if the bank is a subsidiary) is not statistically significant. There appears to be some evidence of a "*flight–to–home*" effect, with banks cutting foreign lending by more than domestic one, according to the OLS estimator. The sum of $SovShock + SovShock \times foreign = -4.7^{***}$ and it precisely estimated, however the difference between the two, the interaction term, is significant only at 10%. According to the GMM estimate, instead, the difference between foreign and domestic lending is small and not statistically significant. In conclusion, the fact that the coefficient of interest, β_1, is still negative and significant even for foreign loans is another indication that $SovShock_{b,t}$ identifies a credit supply channel.

GIIPS and non–GIIPS losses One possible concern with the above estimates, especially those on domestic lending, is that they are the result of weak credit demand for GIIPS banks, whose effect is not completely absorbed by the country–time effects. To alleviate this concern, I split the losses of $SovShock_{b,t}$ between GIIPS ($SovGIIPS_{b,t}$) and non–GIIPS ($SovnonGIIPS_{b,t}$) exposure and I analyze the effect of the two separately for each group of banks, GIIPS and non–GIIPS banks. Table 1.8 and 1.9 show the results for domestic and foreign lending respectively.

First of all, all the coefficients are divided by the standard deviation in each relevant group. This is to ease comparison across the estimates: losses tend to be higher for GIIPS exposure and higher still for GIIPS banks. For example, there is only one bank–year observation with GIIPS losses higher

Table 1.8: Domestic Lending: GIIPS vs non–GIIPS losses

	(1)	(2)	(3)
SovGIIPS/sd	-7.323***		
	(2.74)		
SovnonGIIPS/sd	-1.652		
	(2.48)		
SovGIIPS×GIIPSbanks/sd		-7.617***	-8.531***
		(2.620)	(2.994)
SovGIIPS×nonGIIPSbanks/sd		-4.893***	-4.727***
		(1.824)	(1.775)
SovnonGIIPS×GIIPSbanks/sd			-1.373*
			(0.725)
SovnonGIIPS×nonGIIPSbanks/sd			-0.523
			(3.135)
N	217	217	217
N of banks	89	89	89
R^2	0.634	0.650	0.653

Cluster robust *s.e.* in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note: The dependent variable is the growth rate of domestic loans $\Delta Loans_{b,c,t}$. $SovGIIPS$ and $SovnonGIIPS$ are the bank–specific sovereign losses on GIIPS and non–GIIPS exposure respectively. $GIIPSbanks$ and $nonGIIPSbanks$ are dummies for whether the bank is located in the GIIPS or not. Each coefficient has been divided by the standard deviation in each group, which are: 1.8% for $SovGIIPS$ and 0.2% for $SovnonGIIPS$ in column (1); 2.6% for $SovGIIPS \times GIIPSbanks$ and 0.26% for $SovGIIPS \times nonGIIPSbanks$ in column (2); 0.02% for $SovnonGIIPS \times GIIPSbanks$ and 0.25% for $SovnonGIIPS \times GIIPSbanks$ in column (3).Other balance sheet variables are defined as before. All std.err. have been clustered at the bank level.

than 1% among non–GIIPS banks, so that the interpretation of the coefficient as an elasticity to a 1% loss is an extreme scenario for a non–GIIPS bank. Column (1) in Table 1.8 shows that only losses coming from GIIPS exposure are significantly correlated with domestic lending. It implies that for a one standard deviation shock to GIIPS losses (1.8%) the growth rate of domestic lending declines by 7.3%. Column (2) splits the GIIPS losses between GIIPS and non–GIIPS banks and column (3) adds the same split for non–GIIPS losses too, for a total of four interactions. The key message of column (2) and (3) is that the coefficient on $SovGIIPS \times nonGIIPSbanks$ is also negative and significant: it implies that there is an effect on domestic lending also in countries not under stress. If only the coefficient on $SovGIIPS \times GIIPSbanks$ were significant, then one may worry that the results are all driven by credit demand going down for the most exposed GIIPS banks in GIIPS countries, but this is not the case: there is a supply effect also for non–GIIPS bank exposed to GIIPS debt. In terms of the magnitude, the elasticity is smaller for non–GIIPS banks, but the standard deviation of the shock in that group is ten times smaller.

Table 1.9 repeats the same exercise for foreign lending. Once again, only losses coming from GIIPS exposures have a significant and negative effect on lending: column (1) indicates that a one standard deviation shock (1.6%) decreases lending by 7.7%. Column (2) explores whether this effect is differentiated by whether the *parent* bank is in the GIIPS or not. Differently from the effect on domestic lending, now only GIIPS parent institutions "export" the effect of a local shock to their sovereign on foreign lending through their international subsidiaries, whereas non–GIIPS parent banks do not. But where are the affected subsidiaries of GIIPS banks located? Columns (3)–(5) provide the answer: the effect of the shock is mostly present for subsidiaries in other European countries, in particular in Eastern Europe [20]. Two large Italian banks (UniCredit and Intesa SanPaolo)

[20] I only have a few banks (21) with subsidiaries outside Europe in my sample. The results in column (4) therefore are not surprising in light of the small sample size. Also, given the small number of clusters in this case, the std.err. are White robust std.err

Table 1.9: Foreign Lending: GIIPS vs non–GIIPS losses

	All Foreign (1)	All Foreign (2)	Europe (3)	non–Europe (4)	Eastern Europe (5)
SovGIIPS/sd	-7.471***				
	(2.38)				
SovnonGIIPS/sd	0.635				
	(1.387)				
SovGIIPS× GIIPSparent/sd		-6.624***	-7.157***	-9.47	-5.490**
		(2.139)	(2.582)	(33.5)	(2.093)
SovGIIPS× nonGIIPSparent/sd		-1.188	-1.652	4.590	0.200
		(1.380)	(1.528)	(4.441)	(0.992)
SovnonGIIPS× GIIPSparent/sd		-0.665	-2.951	202.5	1.961
		(6.301)	(6.907)	(145.8)	(5.540)
SovnonGIIPS× nonGIIPSparent/sd		3.360	-0.412	-38.86	11.12
		(8.253)	(8.817)	(126.2)	(6.822)
N	382	382	305	60	182
N of banks	140	140	113	21	68
R^2	0.477	0.478	0.454	0.710	0.673

Cluster robust *s.e.* in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note: The dependent variable is the growth rate of foreign loans $\Delta Loans_{b,c,t}$. $SovGIIPS$ and $SovnonGIIPS$ are the bank–specific sovereign losses on GIIPS and non–GIIPS exposure respectively. $GIIPSbanks$ and $nonGIIPSbanks$ are dummies for whether the bank is located in the GIIPS or not. Each coefficient has been divided by the standard deviation in each group, which are: 1.6% for $SovGIIPS$ and 0.2% for $SovnonGIIPS$ in column (1); 1.4% for $SovGIIPS \times GIIPSbanks$ and 0.12% for $SovGIIPS \times nonGIIPSbanks$; 0.2% for $SovnonGIIPS \times GIIPSbanks$ and 1.2% for $SovnonGIIPS \times GIIPSbanks$ in columns (2)-(5). Other balance sheet variables are defined as before. All std.err. have been clustered at the bank level, but in column (4) where, given the small number of clusters, they are White–robust std.err.

have significant presence in this area, but also Greek, Spanish and Portuguese banks.

Robustness Tests I will now present a series of robustness test on the main result on the baseline regression on domestic loans. All the results are presented for the OLS–FE estimator only, but note that, in most cases, the results hold with the GMM Difference estimator too, both with and without the lagged dependent variable. I will be pointing out any significant departure in specific cases.

Table 1.10: Robustness to Outliers, Credit Demand Controls and Coupon assumptions

	No Greek Banks (1)	$\Delta GDP_{c,t} \times D_c$ (2)	$BLSDem_{c,t} \times D_c$ (3)	Zero Coupon (4)	Par–Zero Average (5)
SovShock	-3.585^{**}	-4.631^{***}	-3.683^{***}		
	(1.796)	(0.956)	(1.364)		
SovShock/std.dev.				-7.581^{***}	-8.170^{***}
				(1.776)	(2.211)
N	206	217	162	217	217
N of banks	83	89	68	89	89
bank FE	yes	yes	yes	yes	yes
year FE	no	yes	yes	no	no
country×year FE	yes	no	no	yes	yes
Bank controls: Tier1(+)***, Profits(+)*** , NPL(-)***, Dep(+), Cash(+)					

Cluster robust *s.e.* in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note: The dependent variable is the growth rate of domestic loans $\Delta Loans_{b,c,t}$. $SovShock$ is the bank–specific sovereign loss. Column (1) excludes Greek banks; column (2)–(3) substitute the country*time fixed–effects with, respectively: GDP growth interacted with country dummies and BLS demand questions (diffusion index, country aggregate) interacted with country dummies. Column (4) uses the zero coupon bond duration for the calculation of the sovereing shock. All std.err. have been clustered at the bank level.

Robustness to outliers, credit demand controls and coupon assumption Table 1.10 tests the robustness of the first set of results. First, I want to make sure that the results are not driven by a few very large outliers. Accordingly, column (1) excludes Greek banks that had the highest losses on sovereign bonds: the results are unchanged and the coefficient is only slightly smaller, -3.6 compared to -4.0.

Columns (2)-(3) verify the robustness of the result to alternative measures of credit demand at the country level. A popular credit demand control in the bank lending channel literature (Altunbas et al. (2009), De Santis and Surico (2013)) is the growth rate of GDP in the country where the bank is lending. Alternatively, the Euro Area Bank Lending Survey (BLS) provides European banks' perceptions on credit demand conditions for the previous three months at a quarterly frequency. The BLS data is available, at the aggregate level [21], for most European countries [22]. I introduce these alternative credit demand controls by interacting either measure with the respective country dummy (columns (4) and (5)). The coefficient is negative and significant in all specifications. The magnitude is very similar to the baseline model with country–time fixed effects [23]. Finally, in column (4) and (5) I modify the coupon bond duration assumption used in the computation of the sovereign shock. Column (4) uses the zero coupon bond duration while (5) averages par and zero coupon bond. Since this alters the entire distribution of the sovereign shock, I divide by the standard deviation to ease comparison with the baseline result in Table 1.6. The coefficient is remarkably similar to the one estimated with the par bond, implying that for a one standard deviation shock using the zero coupon assumption, loan growth decreases by 7.5% vis–a–vis 7.7% with the par bond. The average of the two, that should contain less measurement error than either of the two since it is a better approximation to real sovereign bonds, gives in fact an even larger effect (-8.5%), providing some evidence of an attenuation bias in the other estimates.

Robustness to loan demand homogeneity and simultaneity According to the model in Section 3, I can identify a credit supply effect of sovereign losses only if loan demand is homogenous across banks. Specifically, the loan interest rate elasticity (α_0), that enters into the reduced–form,

[21] Unfortunately, bank–by–bank figures are confidential. Individual BLS demand questions would be a good candidate to control for *bank–specific* credit demand.

[22] The exceptions are non Euro countries such as the UK, Denmark, Norway and Hungary. For Greece and Finland no BLS data exist.

[23] In these robustness tests, the GMM Difference estimator works everywhere but for the BLS demand questions regressions. The coefficient on the sovereign shock in that case is not significant at 5% (p–val 8.9%).

Table 1.11: Robustness to Loan Demand Homogeneity

	$\lambda_{c,t}\times$ size (1)	$SovShock\times$ size (2)	2010Q1 Exposure (3)
SovShock	-4.496^{**}	-4.283^{***}	
	(1.784)	(1.574)	
SovShock$\times$*large*		0.883	
		(0.796)	
SovShock/std.dev.			-3.174^{**}
			1.315
N	217	217	209
N of banks	89	89	84
bank FE	yes	yes	yes
country$\times$year FE	yes	yes	yes

Cluster robust *s.e.* in parentheses
* $p < 0.10$, **$p < 0.05$, ***$p < 0.01$

Note: The dependent variable is the growth rate of domestic loans $\Delta Loans_{b,c,t}$. $SovShock$ is the bank–specific sovereign loss. Column (1) interacts the country–time fixed–effects ($\lambda_{c,t}$) with a dummy $large_b = 1$ if the bank is above the median asset size, 0 otherwise; column (2) interacts $SovShock$ with $large_b = 1$; All std.err. have been clustered at the bank level.

partial effect of the sovereign shock on bank lending (β_1 in the regression), needs to be the same for all banks. Also, the demand shifter elasticity (α_1) is the same across banks *within* the same country–time. In Table 1.11, columns (1) and (2), I relax these assumptions. I do so by interacting both the country–time fixed–effect and the sovereign shock with a dummy $large_b$ equal to one if the bank is above the median size by assets and zero otherwise. The results are largely unchanged [24]. Finally, column (3) uses an extra degree of caution in the construction of the sovereign shock, $SovShock_{b,t}$, by fixing the sovereign exposure at the *pre–crisis* level (March 2010) and letting only the duration and the yield vary over time. The choice of lending to firms and sovereigns are taken simultaneously, thus the one–year lag in the sovereign exposure as defined in (1.1) may not be sufficient to avoid the endogeneity bias. The results are robust to this specification: the coefficient is still negative and significant, although smaller in magnitude.

An IV: Political Ownership One concern with the above estimates is that the amount of

[24]In unreported results, other thresholds for bank size, at the 75^{th} and 25^{th} worked as well.

sovereign bonds in banks' balance sheet is endogenous, so that the sovereign shock would be correlated with unobserved credit conditions. Thus, I can instrument the endogenous part of the sovereign shock, the domestic sovereign exposure over total assets, with the share of bank ownership held by the domestic government in each bank, as measured in De Marco and Macchiavelli (2014).

Table 1.12: IV: Political Ownership in the Panel

	IV: Par (1)	OLS: Par (2)	IV: Zero (3)	OLS: Zero (4)	IV: Parzero (5)	OLS: Parzero (6)
SovShock/std.dev.	-2.491	-3.268**	-3.031***	-3.511***	-2.865**	-3.504***
	(1.527)	(1.496)	(1.043)	(1.076)	(1.175)	(1.224)
N	199	217	199	217	199	217
1^{st} stage F–stat	7.046		16.49		12.45	
bank FE	no	no	no	no	no	no
country×year FE	yes	yes	yes	yes	yes	yes

Robust standard errors in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note: The dependent variable is the growth rate of domestic loans $\Delta Loans_{b,c,t}$. $SovShock$ is the bank–specific sovereign loss. Columns (1)–(2) use the par bond assumption in the calulation of the duration; column(3)–(4) the zero coupon bond assumption and finally columns (5)–(6) take an average of the two. All other bank balance sheet controls are defined as before (not shown). All std.err. are heteroskedasticity robust.

Table 1.12 compares the estimates of the IV and OLS using different duration assumptions. Note that all specifications include the country×year fixed–effect, but not the bank fixed–effect. In fact, since we measure the percentage of political ownership in each bank before the crisis, the IV, $SovPolitical_{b,t} = Duration_{10Y,t} * \Delta yield_{10Y,t} * Political_b$, is time–varying only because the country yield and duration vary over time. Thus, introducing a bank fixed–effect as in the previous specifications does not work with the IV. Column (1)–(2) use the par bond assumption, but the IV estimate is not significant and suffers from a weak instrument problem: the F–stat from the first–stage is below the rule of thumb of 10. Using the zero coupon assumption or the average of zero and par bond instead yields significant results with larger first–stage F–stat. An Hausman test of the equality between the OLS and IV in this case cannot reject the hypothesis that the two coefficients are the same. The estimates imply that for a standard deviation increase in the sovereign shock

(4.2% for zero coupon and 3% for the par–zero average) the growth rate of lending would decline by around 3–3.5%. These numbers are different from those estimated in the above specifications (around 7.5-8%), because I am not controlling for the bank fixed–effect.

An alternative to the bank within transformation is the first differencing estimator that equally removes the bank fixed–effects. Since the the bank–year panel is only for 3 years, I can basically run two separate cross–sectional regression by taking the difference of the growth rate of lending in 2011 against the growth rate of lending in 2010 or 2009. Table 1.13 presents the results.

Table 1.13: IV: Political Ownership in the Cross–Section (differencing)

	2011-2010			2011-2009		
	Par	Zero	Parzero	Par	Zero	Parzero
	(1)	(2)	(3)	(4)	(5)	(6)
SovShock/std.dev	-6.424**	-5.189**	-5.482**	-8.593***	-7.149***	-7.482***
	(3.145)	(2.087)	(2.340)	(3.214)	(2.751)	(2.750)
N (=N of banks)	72	72	72	70	70	70
1^{st} stage F–stat	4.541	14.89	10.48	5.007	13.95	10.79

Cluster robust *s.e.* errors in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note: The dependent variable is the difference in the growth rate of domestic loans between 2011 and 2010 or between 2011 and 2009. *SovShock* is the bank–specific sovereign loss. Columns (1) and (4) use the par bond assumption in the calulation of the duration; columns (2) and (5) the zero coupon bond assumption and finally columns (3) and (6) take an average of the two. All other bank balance sheet controls are defined as before (not shown). All std.err. are clustered at the bank level.

First of all, even in this case the sovereign shock calculated under the par–bond assumption suffers from a weak instruments problem, as the F–stat is below 10. However, both the zero coupon and the average of zero and par bond deliver higher F–statistics and have significant coefficients. The numbers are now much closer to the one estimated with the within–estimator, as expected: a one standard deviation increase in the sovereign shock reduces the growth rate of lending by around 5 to 7%, depending on the specification.

Effect on Syndicated Loans Volume As a further robustness check, I can run the same regression model on the volume of syndicated loans as in Popov and Van Horen (2013). The advantage of using syndicated loan data is that one knows the identity and the location of the borrower, so

Table 1.14: Volume of syndicated loans by country $\log(Loans_{b,c,t}) = \beta_1 SovShock_{b,t} + \eta_b + \lambda_{c,t} + \gamma' X_{b,t-1} + \epsilon_{b,c,t}$

	Country–Borrower FE (1)	Country–Quarter FE (2)	US & EU Borrowers (3)
SovShock/std.dev.	-8.735^{***}	-5.678^{*}	-7.628^{**}
	(2.354)	(3.105)	(3.835)
N	5617	5617	3559
N of banks	74	74	74
bank FE	yes	yes	yes
country–borrower FE	yes	no	no
country×quarter FE	no	yes	yes
Bank controls: Tier1(+), Profits(+), NPL(-), Dep(-), Cash(+)			

Cluster robust *s.e.* in parentheses
* $p < 0.10$, **$p < 0.05$, ***$p < 0.01$

Note: The dependent variable is the volume of syndicated loans of bank b to country c at quarter t. $SovShock$ is the bank–specific sovereign loss; other balance sheet variables are defined as before. Column (1) controls for country–borrower FE; column (2) country×year FE; column (3) country×quarter FE. All std.err. have been clustered at the bank level.

that the country–time fixed–effects is a better control for credit demand. The disadvantage however is that the exact loan breakdown for each lender in the syndicate is not available for the vast majority of loans, so that one needs to create some "artificial" variation. I follow Popov and Van Horen (2013) and divide the loan equally among syndicate members whenever the exact loan shares are not available. The loans are then aggregated at the bank–country–borrower pair at a quarterly frequency. There are a total of 95 country-borrowers, both advanced and emerging markets, for a total of 12,067 loans made by 74 EBA banks over the 2009-2012 period. The results are provided in Table 1.14: column (1) controls only for a country–borrower fixed–effects, column (2) adds an interaction with the quarter dummies: the results imply that, for a one standard deviation increase in the sovereign shock, lending contracts, on average, by 5% to 8%.

1.4.2 Loan Interest Rates

So far, I have shown that European banks with larger losses from sovereign debt tightened their credit supply by reducing aggregate and syndicated lending. However, there is another dimension to credit supply: loan prices or interest rates [25]. If, controlling for credit demand, we see equilibrium interest rates on loans rising, then it must be because of a negative credit supply shock. Since, in the model, banks are assumed to be monopolistically competitive, substituting the equilibrium condition for loans in the downward sloping demand function gives an equilibrium interest rate, r_t^L, as an increasing function of the sovereign shock.

I show that indeed interest rate spreads are on average 40 to 65 bps. higher, depending on the specification, in deals where lenders are hit with a one standard deviation shock to their sovereign portfolio. I am restricting the analysis to banks listed as Lead Arrangers (LA), assuming that these are the relevant lenders with the pricing power in each deal [26]. The sample includes deals with multiple LA, which make up for more than half of the total deals (the median is 2 arrangers per deal as shown in Table 1.2). Therefore, if I were to run the model using the multiple LA sample, significance values would be inflated because of repeated values in the dependent variable. In fact, the all–in drawn spread is the same in each deal even if there are multiple arrangers. To address this concern, I run the model by constructing an "artificial" *average* bank, averaging over balance sheet variables and, especially, the sovereign shock across banks in each syndicated loan. Thus here the sovereign shock is the average shock across lenders (LA) in each deal.

Table 1.15 presents the results. In column (1) I control for country–quarter fixed–effects, whereas in column (2) I exploit the finer disaggregation of the loan–level data and control for country–

[25] A loan has also other non price terms, such as maturity, collateral and debt covenants. However, I do not find any effect of the sovereign shock on these measures. In particular, I do not find evidence that banks with more losses increase maturity of syndicated loans or that debt covenants become tighter (using the covenants strictness measures defined in Murfin (2012)). I do not have good data on the collateral quality.

[26] Admittedly, this assumption may fail if most of the bargaining power in the syndicate is in the hands of the "marginal" participant that is needed to close a deal. To address this concern, in a robustness test I run the model on single–lender deals only.

Table 1.15: Interest Rate Loan Spreads. $Spread_{b,f,t} = \beta_1 SovShock_{b,t} + \eta_b + \lambda_{c,i,t} + \phi' F_{f,t} + \gamma' X_{b,t-1} + \epsilon_{b,f,t}$

	Country–quarter FE (1)	Country–industry–quarter FE (2)	Firm borrower controls (3)
SovShock/std.dev.	40.78^{***}	40.40^{**}	65.41^{**}
	(10.87)	(0.011)	(29.21)
Leverage_ratio			77.40^{**}
			(33.82)
Log(Assets)			-40.50^{***}
			(11.41)
EBITDA/Sales			-12.29
			(16.51)
Investment/Assets			-4.160^{***}
			(1.158)
N	5147	5147	949
bank FE	yes	yes	yes
country×quarter FE	yes	no	no
country×industry ×quarter FE	no	yes	yes

p-values in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note: The dependent variable is the (log of) the all-in drawn spread on loans made by Lead Arranger b to firm f at quarter t. $SovShock$ is the bank–specific sovereign loss at a quarterly frequency, divided by its standard deviaion; other balance sheet variables are defined as before. Column (1) controls for country–borrower FE; column (2) country×year FE; column (3) country×quarter FE. All std.err. have been clustered at the bank level.

industry–quarter fixed–effects [27]. Basically, in column (2) I am comparing the interest rate charged by an (*average*) bank hit with a one standard deviation shock and a bank not hit by the shock when they lend to the corporate borrowers in the same sector, country and quarter. In terms of the model in Section 3, I am allowing for the demand shifter to be not just country–time specific ($\lambda_{c,t}$), but country–industry–time specific ($\lambda_{c,i,t}$). The results suggest that the interest rate loans made by banks hit with a one–standard deviation shock [28] are 40 bps (one quarter of the standard deviation of interest rates) higher than banks with no shock. Furthermore, column (3) uses the DealScan–Compustat sample to control for firm–level characteristics. The effect of the shock is still positive and significant and it implies an even larger effect of 65 bps increase in spreads for a one standard deviation increase in the shock. Other firm characteristics have the expected sign: more levered and smaller firms pay higher interest rate spreads [29].

The regressions in Table 1.15 are robust to changing the various assumptions underlying the construction of the sample. In particular, to ease concerns that the results are driven by outliers that skew the shock distribution for the "artificial" *average* bank, in Table 1.16 column (1) I restrict the sample to the largest LA by total assets in each deal. These banks are mostly global banks with smaller shocks and they are more likely to be those with the most pricing power in each deal: the effect is just slightly smaller, 30 vs 40 bps higher. Furthermore, if one worries that the assumption of assigning the pricing power to LA is not accurate, column (2) analyzes single–lender deal only, nearly half of which are listed as non–LA in DealScan: the effect is larger than in the baseline specification (63 bps.), although significant at around 5% only. In the rest of the columns in Table

[27] I cannot control for firm– or firm–time fixed–effects because I do not observe many firms borrowing in more than one deal in my sample. In fact, the average maturity of syndicated loans is 5 years and I am focusing on a 3 year window, 2010-2012

[28] The standard deviation of the sovereign shock in this sample is about 0.2% or 20 bps in terms of losses over total assets. No bank has *quarterly* losses of 1% of total assets, so I find it more realistic to provide the results normalizing by the standard deviation

[29] The result is not driven by the inclusion of firms' covariates. In an unreported robustness test, running the regression on the Dealscan–Compustat sample without including borrowers' balance sheet characteristics yields the same result.

1.16 I run other robustness cheks. Column (3) and (4) distinguish between credit lines or term loans, keeping the average bank assumption. The effect of the shock appears not to be significant for credit lines, but it is significant and even stronger for term loans. Finally Column (5) adds two loan characteristics: the (log of) maturity and the (log of) loan amount. As expected, larger loans and those with shorter maturities have higher interest rate spreads, but the effect of the sovereign shock is still positive and significant.

Table 1.16: Interest Rate Spreads Robustness Tests

	Largest LA (1)	Single Lender (2)	Credit Lines (3)	Term Loans (4)	Loan controls (5)
SovShock/std.dev.	30.31^{**}	63.70^{*}	15.57	76.27^{***}	30.89^{**}
	(14.57)	(32.25)	(19.64)	(23.38)	(13.13)
Log(Loan Size)					-26.37^{***}
					(4.021)
Log(Maturity)					40.40^{***}
					(4.338)
N	5147	1372	2346	2377	5074
bank FE	yes	yes	yes	yes	yes
country×industry ×quarter FE	yes	yes	yes	yes	yes

Cluster robust *s.e.* in parentheses

* $p < 0.10$, **$p < 0.05$, ***$p < 0.01$

In conclusion, I have shown that sovereign losses matter for credit supply not only because they reduced the growth rate of credit, but also because they increase the interest rate spreads charged on syndicated loans. The effect is not driven by sector specific credit demand, it does not depend on assumptions on lenders pricing power, loan characteristics and, most importantly, holds also after I control for borrower specific characteristics.

1.5 The channels

I have established that the sovereign debt crisis has had a negative effect on the supply of loans through its effect on banks' balance sheets. In this section, I will explore two main hypotheses as to why sovereign exposures matter for credit supply: the *capital channel* and the *funding channel*.

According to the first, banks that need to recapitalize during the crisis period prefer to do so by shedding assets (loans) than by raising new equity. The regulatory capital target in Europe, the Core–Tier1 ratio (CT1), is defined as common equity, including government support measures, over Risk–Weighted Assets (RWA). Notably, government bonds receive a 0% risk weight in the calculation of RWA. When a negative equity shock (losses on the sovereign bond portfolio in this case) occurs, banks may go below the minimum level of regulatory capital. They can get back to the target ratio by either raising equity or by reducing risky assets, especially in the loan portfolio. However, we know that equity is a relatively costly source of finance (Myers and Majluf (1984)), so that a bank may be reluctant to issue new shares, especially at a time of low stock prices [30]. If this is the case, and capital constraints are *binding*, cutting off loan supply, thus reducing RWA, seems the only viable alternative to increase the capital ratio [31]. Therefore, this channel may not be at work if capital constraints are not *binding* and the *potential* sovereign losses are not realized on bank books. It is difficult to ascertain whether capital constraints are binding in practice, because theoretical models (Repullo and Suarez (2009)) predict that banks would hold capital buffers well in excess of the minimum requirement, but could still find the constraint binding in their optimization problem. As I discuss at the end of this sections, there are reasons to believe that capital constraints were not binding over this period. Regarding the accounting value of sovereign losses, it is true that

[30] Other ways to increase equity without issuing new shares include: increase retained earnings (difficult to do in the short term), debt–to–equity and hybrid shares conversion (widely used according to EBA and BIS (2012) reports).

[31] According to EBA and BIS Quarterly reports (2012), another way to reduce RWA without asset shedding is to change the risk weights used in internal models. Apparently, these changes were pre–agreed with regulators and they were used extensively during the sovereign debt crisis. I take this into account normalizing equity by total assets as well as RWA.

on average 40% of banks' sovereigns are in the HTM banking book, where they are not marked–to–market. However, according to the EBA September 2011 recommendation, capital had been assessed net of valuation losses on the sovereign portfolio. Banks had to mark–to–market their whole sovereign portfolio, including the HTM banking book. Thus, in principle, these losses had to appear on banks' books and banks had to put up capital against it [32].

The *funding channel*, on the other hand, suggests that losses on sovereign bonds matter for credit supply because they impair banks' ability to refinance on the wholesale market, especially in the secured interbank market (repos). Government bonds are the preferred source of collateral used for interbank repos, where the size of the haircut, the repo rate and the maturity depend on the perceived risk of the collateral [33]. When tensions on sovereign markets reached high levels in 2010–2012, banks lacked an important source of funding and this could have reduced the capacity to provide credit to the real economy (Gonzalez–Paramo (2011)). For example, Figure 1.5, which I obtained from the International Capital Markets Association (ICMA) survey, shows that by June 2012 virtually no European bank could use Greek, Irish or Portuguese debt as collateral for interbank repos [34]. My measure of the sovereign shock represents a proxy for the average quality of the collateral that banks can post on the interbank market. Admittedly, since the ECB eligibility criteria and haircuts for collateral have been less stringent than market ones throughout the debt crisis, the total effect on the funding channel may be ambiguous. In fact, if banks could not refinance on the open market, they could always resort to the ECB marginal lending facility that switched to a fixed rate auction with full allotment in late 2008 or, especially, participate in the 3–year longer–term refinancing

[32] Sovereign exposures in the AFS book are measured at fair–value and exposures in the HTM are valued "in a conservative fashion, reflecting market prices as of 30 September 2011" (EBA, Methodology for Recap Exercise 2011).

[33] The European repo market is fundamentally different from the US one. Mostly it is bilateral and traded via a central clearinghouse. The haircuts are set by the clearinghouse and not by market participants, however they are still affected by market developments (see Boissel et al.(2014)).

[34] The absence of "bad" banks with "bad" collateral has led some (Mancini et al. (2013)) to conclude that the repo market was very resilient, even in crises periods. Boissel et al. (2014), using micro data on bilateral repos broken down by country, show instead that average repo rates in GIIPS countries increase with the sovereign CDS.

operations (LTRO) in December 2011 and February 2012 that injected a total of €1.1 tn. in the banking system (Drechsler et al. (2012)).

Table 1.17: The Capital and the Funding Channel

	Capital Channel: Regulatory (1)	Capital Channel: Leverage (2)	Funding Channel (3)	Funding Channel: 2008Q4 (4)	Both Channels (5)
$lowTier1/RWA$	-0.00265				-0.0401
	(0.019)				(0.033)
$lowTier1/TA$		0.00204			
		(0.051)			
$highShortTermFund$			-0.0246		-0.0194
			(0.042)		(0.042)
SovShock	-4.049^{**}	-3.504^{*}	-3.538^{**}	-3.521^{**}	-3.240^{*}
	(1.612)	(1.81)	(1.681)	(1.714)	(1.667)
SovShock×	0.157				0.624
$lowTier1/RWA$	(0.509)				(0.438)
SovShock×		-1.243			
$lowTier1/TA$		(1.597)			
SovShock×			-1.707^{***}	-2.938^{**}	-2.106^{***}
$highShortTermFund$			(0.788)	(1.276)	(0.749)
Shock+Shock×	-3.892^{**}				-2.615
$lowTier1/RWA$	(1.79)				(1.799)
Shock+Shock×		-4.743^{***}			
$lowLeverageRatio = 1$		(1.719)			
Shock+Shock×			-5.235^{***}	-6.459^{***}	-5.347^{***}
$highShortTermFund = 1$			(1.387)	(1.648)	(1.347)
N	217	216	216	216	216
N of clusters	89	89	89	89	89
bank fixed effects	yes	yes	yes	yes	yes
country*time fixed effects	yes	yes	yes	yes	yes
Bank controls: Tier1(+)***, Profits(+)*** , NPL(-)***, Dep(-), Cash(+)					

Cluster robust $s.e.$ in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note: The dependent variable is the growth rate of domestic loans $\Delta Loans_{b,c,t}$. $lowTier1\_nogovhelp_{b,t}$ takes value 1 if the bank is below the 25^{th} pct. of the level of Tier1 ratio in year t, 0 otherwise; $lowLeverageRatio_{b,t}$ is equal to 1 if the bank is below the 25^{th} pct. of the level of the leverage ratio (CommonEquity/Total Assets) in year t, 0 otherwise; $highShortTermFund_{b,t}$ takes value 1 if the bank is above the 75^{th} pct. in short–term funding over total funding in year t, 0 otherwise; other balance sheet variables (Tier1, Deposits/Assets, PreTaxProfits/Assets, Cash/Assets) are defined as before. Standard errors are clustered at the bank–level.

Table 1.17 explores the capital and the funding channel in greater detail. The dependent variable in all regression is the growth rate of domestic loans, as in the baseline results. Column (1) tests for the presence of the capital channel by interacting the sovereign shock with a dummy variable,

$lowTier1/RWA_{b,t}$, that takes value one if the bank, in each year, has a low regulatory capital (below the 25^{th} pct.) and zero otherwise. The slope coefficient on the sovereign shock is significant and negative for all banks, regardless of the level of regulatory capital, but there appears to be no additional negative effect for poorly capitalized banks. If anything, in the subtable I show that the total impact of a 1% increase in the sovereign losses over assets ratio on loan growth seems to be smaller in absolute value for low capitalized banks (-3.9%) than for the other banks (-4.7%), but the difference is not statistically significant. Using other thresholds to define the dummy $lowTier1_{b,t}$ (either higher (50^{th}), lower (10^{th}) or fixed thresholds imposed by the EBA (5% in 2010 9% in 2011)) give similar results (not shown in the table).

Column (2) explores the capital channel using another definition of bank capital, the actual *leverage ratio*, defined as Tier1 equity over Total Assets (not RWA), as it appears in Basel III rules. In fact, even though sovereign losses may not have an impact on regulatory equity, the bank may still decide to de–leverage so to ease investors' concerns on the solvency of the bank. In fact, even though market participants may not care about regulatory equity and the results of stress tests, they are still interested in the overall solvency of the institution. Accordingly, the dummy used in the interaction in column (2), $lowTier1/TA_{b,t}$, is equal to one if the bank is below the 25^{th} pct. of the distribution of the leverage ratio in each year and zero otherwise. The interaction term is not statistically significant and the baseline effect for the banks with a leverage ratio above the first quartile is significant only at around 5%. However, if I compute the total effect of the sovereign shock on credit growth for banks that have a low leverage ratio, I find that it is larger (-4.73%) than the total effect for banks with low regulatory capital ratio (-3.8%) and it is significant at the 1% level. This suggests that indeed some of the de–leveraging happened because of banks' sovereign losses, although not as a consequence of regulatory action.

Column (3) turns to the funding channel. Here the interaction is with a dummy $highShort-$

$-TermFund_{b,t}$ that takes value 1 if the bank is above the 75^{th} pct. in short–term funding over total funding [35]. The interaction term is negative and significant, implying that for banks with a higher dependence on short term funding there is an additional negative effect of sovereign losses on bank lending. The total effect of the sovereign shock for these banks is a decrease in the growth rate of loans of 5.089% compared to 3.203% for other banks. It is also very precisely estimated. Note that this effect is present only for banks *highly* dependent on this source of funding: other thresholds at 90^{th} pct. work, but not at the 50^{th}.

A possible concern with the regression in column (3) is that the dependence on short term funding is endogenous, especially during a crisis: distressed banks could be forced to substitute long–term, stable source of funding (such as customer deposits) with short term debt. So it could be that the dummy $highShortTermFund_{b,t}$ is picking up solvency rather than funding liquidity concerns. One way to address the issue is using the dependence on short term funding at the beginning of the sample, at the end of 2008, to see if banks that "normally" fund themselves with short–term debt have been differentially impacted by the sovereign debt crisis. Therefore, column (4) defines $highShortTermFund_{b,t}$ to be the dependence on short term funding before the sovereign debt crisis, at the beginning of 2009 (2008Q4). The dummy itself, not being time varying, cannot be included in the fixed–effect regression. I still find an additional negative kick for banks highly dependent on short term funding, the total effect is even larger than before, implying a decrease in domestic loan growth of around 6.5%.

Finally, column (5) tests the joint hypothesis that both the regulatory capital and funding channels are working at the same time. Similar results, not shown, apply with the leverage ratio as a measure for the capital channel. It appears that the funding channel largely dominates the capital channel.

[35] Bankscope provides a variable called *Other Deposits and Short–Term Funding* that captures all short term funding not classifiable as customer deposits. This includes interbank repos, but also short term certificates of deposits and all non depository sources of funding. So it is only an imperfect measure of interbank funding, but it should nonetheless capture the extent to which a bank is exposed to a short–term funding shock.

In fact, the estimate of the sovereign shock for the low capitalized banks is much attenuated (-2.6%) and less precisely estimated. On the other hand, the effect for those highly dependent of short term funding is more negative (-5.3%) and significant.

In conclusion, the data seem to support the hypothesis that high sovereign debt exposure to risky sovereign debt affected banks' cost of funding rather than the cost of capital, consistent with the simple theoretical model provided in Section 3. I interpret the absence of the capital channel as a sign that capital constraints were not binding over this period. In fact, the *forbearance* by the EBA in enforcing the capital requirements may be responsible for this: banks were given plenty of leeway into how they could address regulatory capital shortfalls. According to EBA and BIS (2012) reports, almost a third (28%) of the aggregate shortfall in capital by EBA banks could be fulfilled with debt–to–equity and other hybrid shares conversion, rather than by issues of new equity. Notably, Banco Santander of Spain was allowed a €6.83 bil. debt–to–equity conversion vis–a–vis a capital shortfall of €15.3 bil. Another 10% would come from changes to internal models to calculate risk–weights, which were pre–approved by regulators. This type of regulatory (in–)action is perfectly understandable given that risk based capital requirements tend to be pro–cyclical (they rise during recessions), regulators may be reluctant to impose additional capital buffers that would just exacerbate the crisis. On the other hand, *market discipline* would make funding problems unavoidable for banks. If participants in the interbank market believe that the government bonds posted as collateral by a bank are not of sufficient quality or other lenders in general perceive the bank as risky, they may reduce the amount of money they lend to that bank (*i.e.* an increase in the haircut) or increase the repo rate. There would not be any forbearance on part of other market participants and the bank, unable to borrow on the market, has to cut loan supply .

1.5.1 Bank–specific Cost of Funding: Bond Issuance

In the previous section I have provided evidence that sovereign losses matter for banks' cost of funding. Government bonds are special in this respect because they provide a source of collateral for secured interbank transactions (repos). Ideally, in order to test the hypothesis that sovereign losses increased banks' cost of funding, I would look at bank specific repo rates, but unfortunately these data are not available at the bank level and are anonymized (Boissel et al. (2014)). However, the results above are consistent with a broader view of the funding channel. In fact, banks with higher exposure to risky sovereign debt are perceived as more risky by all types of lenders, not only other banks. Therefore, it could be that other sources of funding are also affected by sovereign exposures. One such source for which data are available are banks' bonds issuance from SDC Platinum.

I focus on fixed–rate [36] coupon bonds issued by European banks and their subsidiaries. There is a total of 2,343 bonds with coupon rate data issued by EBA banks or their subsidiaries between 2009 and 2012 available on SDC Platinum. There are 110 different issuers, 49 of which are subsidiaries of 61 EBA banks. A third of the sample (787 bonds) comes from banks in the UK, another half (1,062 bonds) from Dutch, German, French and Italian banks, while non–European banks affiliated to banks in the EBA sample only issued 186 bonds. Table 1.18 contains some summary statistics for the SDC sample. The bonds issues are quite large, with an average of \$373 bil., they are quite skewed towards the right (the median is \$63 mil.) and they have an average coupon rate of 3.7%. The average maturity, which is also the median and the mode, is five years, with very few bonds (27 issues) with maturities shorter than one year. Finally, the ratings, which are available for 75% of the bonds, are very high: the top 25% are of the highest quality (Aaa), but even the bottom 25% is of a high–quality grade (Aa3).

It is important to keep in mind that this source of bank funding is not short–term, as the average

[36] Floating or Indexed bonds do not have a coupon rate

Table 1.18: SDC Platinum Bonds Issuance Summary Statistics

Variable	Obs	Mean	Std.Dev.	Min	25^{th}	50^{th}	75^{th}	Max
Coupon (%)	2343	3.73	2.14	0.01	2.12	3.5	5	11.25
Maturity (years)	2327	4.79	2.38	0	3	5	6	10
Principal ($ mil.)	2343	373.4	648.3	0.046	12.9	62.9	399.4	658
Moody's rating	1750			Ba3	Aa3	Aa2	Aaa	Aaa

bond maturity is around five years. Thus it is very different from repos, which are mostly overnight or with maturity less than a week (ECB (2012)). Nonetheless it could still be true that banks with higher sovereign losses are perceived as more risky and they need to offer higher coupon rates for medium and long–term bonds. I test this hypothesis below.

The empirical specification I employ in this section is very similar to the one used so far in the paper, but with bond coupon rates as the dependent variable. Specifically:

$$Coupon_{b,k,t} = \beta_1 SovExp_{b,t} + \gamma' X_{b,t-4} + \alpha' X_{k,t} + \eta_b + \lambda_{b,c,t} + \epsilon_{b,k,t}$$

where $Coupon_{b,k,t}$ is the coupon rate of bond k issued by bank b (which could be an EBA bank or a subsidiary) in quarter t; $SovExp_{b,t}$ is the sovereign shock for an EBA bank b measured on a quarterly basis, as in the analysis for syndicated loans. Therefore, even if the bond is issued by a subsidiary the sovereign shock is measured at the parent bank level, as in the analysis of foreign loans. $X_{b,t-4}$ are group b characteristics at the beginning of the *year* and $X_{k,t}$ are bond–specific characteristics, such as maturity and rating. I also control for $\lambda_{b,c,t}$, a set of country–quarter (or year) fixed–effects for the country of the parent bank (or issuer, either parent or subsidiary) and a set of bank (*i.e* issuer) fixed–effects, η_b. Table 1.19 presents the results.

An increase in the sovereign shock by one standard deviation (20 bps.) increases the coupon rate by 25 to 50 bps., depending on the specification. All specifications include bank controls, issuer and maturity (in years) fixed–effects, but use a different set of country–time fixed effects. Columns (1) and (2) have country–year effects, either for the country of the parent bank or for the country of the issuing bank: the coefficient implies that for a 20bps. increase in the sovereign shock, the coupon rate increases by 20–25bps, which is equivalent to one tenth of a standard deviation of coupon rates. Column (3) uses country–quarter fixed–effects (for the country of the parent bank) and the coefficient is quite stable, while in column (4), which also includes Moody's ratings, the effect doubles to 52 bps. for a one standard deviation of the sovereign shock. The sign of the coefficient on Moody's ratings is positive, as expected: on average, a decrease in the quality of the bond by one notch implies an increase in the coupon rate by 30bps. Finally, column (5) includes a set of country–*quarter* fixed–effects for the countries where the issuers are located, either a subsidiary or a parent bank. The coefficient is no longer significant. This may be the result of limited variability in the independent variable across countries and quarters, since five countries (UK, Germany, Netherlands, France and Italy) make up 80% of the sample.

1.6 Conclusions

In this paper, I have shown that the sovereign debt crisis has had a negative real effect on credit supply through its impact on banks' balance sheets. Using bank–by–bank exposure data to sovereign debt, I calculate the exact sovereign losses in banks' portfolio and I use them as an explanatory variable for the growth rate of loans and loan interest rate spreads. The results suggest that banks hit by a large sovereign shock (a one standard deviation increase) had a growth rate of domestic loans around 7.7% lower than a bank not hit by the shock. The results are robust to the elimination of outliers (Greek banks), to differences in unobservable borrowers' characteristics between large and small banks, to the exclusion of the home sovereign exposure and to the assumption used to compute the duration (zero–coupon or par bond). To provide conclusive evidence that the sovereign debt crisis represented a negative credit supply shock, I have also shown that global European banks reduced lending abroad, through their international subsidiaries especially in Eastern Europe, by the same amount as domestic loans. Moreover, I find that for a one standard deviation increase in sovereign losses-over-total assets (15–20bps.), banks charge interest rate spreads 40 to 65 bps. higher, even after controlling for industry unobserved heterogeneity and corporate borrower characteristics.

I also attempt to shed some light on the mechanisms as to why sovereign losses matter for bank lending. I find evidence for a *funding channel* over a *capital channel*: sovereign losses affect disproportionately more the growth rate of credit for those banks with a higher share of short term funding rather than those with low level of capitalization. I interpret the results as *forbeareance* from the European regulator (EBA) in enforcing capital requirements in a time of crisis, but *market discipline* from market participants.

Appendix

Table 1.19: Fixed Coupon Rates and the Sovereign Shock

	(1)	(2)	(3)	(4)	(5)
SovShock/std.dev.	24.66***	21.62***	30.35***	52.39**	25.19
	(7.664)	(7.159)	(11.43)	(20.54)	(22.42)
Moody's				0.319***	0.280**
				(0.118)	(0.126)
Bank Controls: Tier1(–), Profits(+), NPL(–), Dep(–), Cash(–)**					
Issuer FE	yes	yes	yes	yes	yes
Years of maturity FE	yes	yes	yes	yes	yes
Countryparent–year FE	yes	no	no	no	no
Countryissuer–year FE	no	yes	no	no	no
Countryparent–quarter FE	no	no	yes	yes	no
Countryissuer–quarter FE	no	no	no	no	yes
N	1876	1876	1876	1391	1391
N of issuers	105	105	105	90	90

Standard errors in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note: the dependent variable is the coupon rate of fixed–coupon bonds. All specifications include a bank (issuer) fixed–effect, a set of dummies for the maturity of the bond (in years) and bank controls. Column (1) includes a country–year FE for the country of the parent bank; column (2) a country–year effect for the country of the issuing bank (either a subs or the parent); column (3)–(4) a country–quarter effect for the country of the parent bank and finally column (5) a country–quarter effect for the country of the issuing bank. Moody's ratings are in a scale from 1 (Aaa) to 5 (Ba3). All std.err. have been clustered at the bank (issuer) level.

Table 1.20: Political banks and Performance. Dependent variable: Political ownership (%)

	(1)	(2)	(3)	(4)
	2010	2010	2011	2011
$NPL/TA_{b,t}$	0.0171*	0.0233	0.00766	0.0180*
	(0.00945)	(0.0147)	(0.00743)	(0.0105)
$Profits/TA_{b,t}$	-3.237*	-2.758	2.904*	0.723
Assets	(1.791)	(3.488)	(1.512)	(2.626)
Log($Assets$)	-0.0485**	-0.0567*	-0.0694***	-0.0694
	(0.0185)	(0.0286)	(0.0237)	(0.0466)
Cash/Assets	-6.909***	-6.584***	-1.625	-0.231
	(1.640)	(2.458)	(1.204)	(1.634)
Tier	-0.122	-0.873	-0.694	-1.369
	(0.826)	(1.114)	(0.836)	(1.287)
Dep/Assets	-0.0515	-0.0691	-0.325	-0.345
	(0.322)	(0.468)	(0.327)	(0.418)
N	87	87	78	78
Country fixed–effects	no	yes	no	yes

White standard errors in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 1.21: Political Banks and Credit Portfolio Breakdown

	Corporate (1)	Retail (2)	RetailSME (3)	RetailResMtg (4)	CRE (5)	Institutions (6)
$Political_b$	-0.0440	-0.120*	-0.0224*	-0.0892	-0.0267*	0.116**
	(0.0475)	(0.0624)	(0.0131)	(0.0615)	(0.0142)	(0.0437)
$\log(TA_{b,t-1})$	0.00895	-0.0225	-0.00947*	-0.0171	-0.00166	-0.00381
	(0.0121)	(0.0151)	(0.00527)	(0.0168)	(0.00382)	(0.00761)
$Tier1/RWA_{b,t-1}$	-1.024*	0.335	0.251	-0.0726	-0.203	0.398
	(0.535)	(0.621)	(0.442)	(0.653)	(0.233)	(0.449)
$Dep/TA_{b,t-1}$	-0.0723	0.418***	0.0250	0.287**	-0.00146	-0.281***
	(0.107)	(0.120)	(0.0702)	(0.135)	(0.0418)	(0.104)
$Cash/TA_{b,t-1}$	0.669	1.341	-0.404	1.005	-0.551*	-1.062
	(0.876)	(1.179)	(0.425)	(1.174)	(0.317)	(0.726)
$NPL/TA_{b,t-1}$	0.582	-2.181**	0.539	-2.538***	0.259	-0.645
	(0.764)	(0.931)	(0.414)	(0.895)	(0.291)	(0.613)
$Prof/TA_{b,t-1}$	5.313***	-0.854	1.347	-2.539	-0.628	0.220
	(1.791)	(2.160)	(1.527)	(2.416)	(0.772)	(1.422)
N	78	78	78	78	78	78

Robust standard errors in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: this table shows the differences in the credit portfolio between government–owned and private banks using EBA data on Exposures at Default (EAD) from the 2011 Stress Test, referring to December 2010. The dependent variables are the shares of the total EAD to all countries in the following portfolios, as classified by the EBA: Corporate excluding Commercial Real Estate, Retail excluding Commercial Real Estate, Retail of which Small–Medium–Enterprises (SME), Commercial Real Estate, Institutions (public entities).

Table 1.22: Political Banks and Credit to Institutions

	Home Institutions (1)	Foreign Institutions (2)	Home Institutions (3)	DomSov /TotSov (4)
$Political_b$	0.0995***	0.0160		
	(0.0323)	(0.0215)		
$Political_b \times$non–GIIPSbanks			0.123***	0.209***
			(0.0365)	(0.0695)
$Political_b \times$GIIPSbanks			0.0418*	0.357***
			(0.0235)	(0.107)
N	78	78	78	78
Other Bank Controls: log(TA),Tier1/RWA,Dep/TA,Cash/TA,NPL/TA,Prof/TA				

Robust standard errors in parentheses

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: this table shows the differences in the Institutions (public entities) credit portfolio between government–owned and private banks using EBA data on Exposures at Default (EAD) from the 2011 Stress Test, referring to December 2010. The dependent variables are the share in lending to domestic public institutions over the total (column (1) and (3)), the share to foreign public institutions (column (2)). Finally, column (4) has the share in domestic sovereign bonds over total sovereign bonds.

A Simple Model of Bank Lending This section describes a simple model of bank lending. The purpose of the model is to help understanding the identification assumptions underlying the empirical strategy. In particular, it spells out the assumptions on loan demand under which, as a result of sovereign losses, a credit supply shock can be identified using banks' balance sheet data. The baseline model is a modified version of Stein (1998)'s paper on banking with adverse selection and it is similar to Ehrmann et al. (2003).

Bank b needs to satisfy the following balance sheet constraint at time t: $L_{b,t}+S_{b,t} = E_{b,t}+B_{b,t}+D_t$. On the asset side, $L_{b,t}$ is loans and $S_{b,t}$ represents (risky) sovereign bonds. For simplicity, banks hold no other security. These assets are funded through equity, $E_{b,t}$, short–term, interbank funding $B_{b,t}$ and customer deposits D_t. Deposits are exogenous. They pay zero interest and are demanded by households as a mean of payment. Their demand is given by $D_t = c - \delta r_t^f$ with $c, \delta > 0$ and r_t^f the risk free rate. Bank capital, under the Basel II regulation, is determined as a fraction of risky assets (loans): $E_{b,t} = \kappa L_{b,t}$ with $\kappa < 1$. Thus the balance sheet can be conveniently rewritten as:

$$B_{b,t} = (1-\kappa)L_{b,t} + S_{b,t} + c - \delta r_t^f \tag{1.4}$$

The interbank funding rate is:

$$r_{b,t}^B = r_t^f + \mu_1 \mathbb{X}(S_{b,t-1}) + \mu_2 f(X_{b,t-1}) \quad \text{with} \quad \partial x/\partial S > 0, \mu_1 > 0, \mu_2 >> 0 \tag{1.5}$$

where $\mathbb{X}$ is an increasing function of the level of *lagged* sovereign exposure and f is a function of other *predetermined* bank characteristics ($X_{b,t-1}$ is a $k \times 1$ vector). μ_1 is a positive constant and μ_2 is a $k \times 1$ vector of positive constants. Thus, if the bank is more heavily exposed to risky sovereign debt $S_{b,t}$, it will face a higher cost of funding for short–term sources of funds. The partial derivatives of the function $f(\cdot)$ with respect to the components of the vector $X_{b,t-1}$ are negative if the element of the vector is a "good" bank characteristic, positive otherwise [37]. These characteristics are not explicitly modeled here. One should think about these as endogenous, but predetermined variables that determined the bank's cost of funding. Likewise, sovereign bonds are an endogenous choice variable, but they are predetermined in determining the bank funding rate for short term funds. I acknowledge this fact in the empirical strategy by estimating the model through the dynamic panel GMM estimator (Arellano and Bond (1991)). Additionally, I can address the endogeneity of sovereign exposure via instrumental variable techniques. I use the percentage of political ownership in each bank as an instrument for its domestic sovereign exposure. Section 4 explains the IV strategy thoroughly.

Thus, in the model, the sovereign shock matters for banks because it increases the cost of funding, but it does not directly affect equity. Any negative equity shock is ruled out because the amount equity is simply tied to the level of risky assets (loans) through the Tier 1 capital ratio. The data seem to support this assumption, as there is evidence that banks with higher dependence on short term funding, and not undercapitalized banks, were more negatively affected by the sovereign shock. Section 7 analyzes this aspect in greater detail.

Banks are monopolistically competitive in the loan market and they all face a downward sloping loan demand when they lend in country c:

$$L^D_{b,c,t} = -\alpha_0 r^L_{b,t} + \alpha_{1,c,t}\lambda_{c,t} \quad \text{with} \quad \alpha_0, \alpha_{1,c,t} > 0 \tag{1.6}$$

where $\lambda_{c,t}$ is an aggregate demand shifter in country c at time t and $r^L_{b,t}$ is the loan interest rate charged to borrowers, α_0 is the loan interest rate elasticity and $\alpha_{1,c,t}$ is the impact of the country demand shifter which is allowed to vary over country and time. Notice that c is not necessarily

[37] In the empirical part, the "good" characteristics are going to be capitalization, profitability and liquidity while the "bad" one is the average quality of the loan portfolio

the country where the bank is headquartered, as banks can lend internationally through their subsidiaries. However, in the empirical section, I assume that no bank directly lends to more than one country: the international subsidiaries are part of the group, but are independently managed.

There are several assumptions behind this loan demand schedule. First of all, since it contains only the loan rate, it implicitly assumes that substitution with other forms of finance is impossible. This may be extreme, but it is nonetheless a good approximation for many corporate borrowers in Europe for which bank funding is the predominant form of credit. Bank debt over total external financing for non financial firms is, on average, well above 80% in most European countries (see Altomonte et al. (2011)). Second, it assumes that loan demand is homogeneous across banks *within the same country*. In fact, the demand shifter is at the country–aggregate level and it is not bank–specific. Its impact ($\alpha_{1,c,t}$) is the same for all banks within a country. Also, notice that the interest rate elasticity (α_0) is the same for all banks. This rules out, for instance, that borrowers of large and small banks have different interest rate sensitivities. I can somewhat relax this set of assumptions in the empirical exercise. For example, I allow the elasticity $\alpha_{1,c,t}$ and α_0 to vary between large and small banks. When analyzing loan interest rates, the demand shifter is not only country specific, but country *and* sector specific.

Banks maximize the future discounted value of dividends (see Appendix for detail) by choosing sequences of loans $L_{b,t}$ and sovereign holdings $S_{b,t}$:

$$\max_{S_{b,t},L_{b,t}} \mathbb{E}_t \sum_{i=0}^{\infty} \beta^i (r^L_{b,t-1+i} L_{b,t-1+i} + r^S_{t-1+i} S_{b,t-1+i} - r^B_{b,t-1+i} B_{b,t-1+i} - \kappa L_{b,t+i} - \phi(L_{b,t+i}))$$

s.to (1.5) and (1.6). Note that B_t is determined residually from (1.4) once $L_{b,t}$ and $S_{b,t}$ are chosen. r^S is the rate of return on risky sovereign (exogenous to the bank) and $\phi(L_{b,t})$ represent costs associated to banking activities, such as evaluation of credit rating of the customer, administering and monitoring the loan. I assume a quadratic cost of servicing the loans plus a bank–specific component: $\phi(L_{b,t}) = b_0/2L^2_{b,t} + \eta_b L_{b,t}$.

Imposing loan market clearing and substituting in (1.4) and (1.5), the FOC for loans and sovereigns are:

$$
\begin{aligned}
(L): \quad & L_{b,t} = \frac{1}{2\beta + b_0\alpha_0}(-(1-\kappa)\alpha_0\mu_1\mathbb{X}(S_{b,t-1}) - (1-\kappa)\alpha_0\mu_2' f(X_{b,t-1}) + \alpha_{1,c,t}\lambda_{c,t} \\
& \qquad - (1-\kappa)\beta\alpha_0 r^f_t - \alpha_0(\eta_b - \kappa)) \\
(S): \quad & r^s_t - \frac{\partial x}{\partial S_{b,t}}((1-\kappa)L_{b,t} + S_{b,t})) - (r^f_t + \mu_1\mathbb{X}(S_{b,t}) + \mu_2' f(X_{b,t})) = 0
\end{aligned}
$$

The main empirical specification is a modified version of the first FOC [38]. The level of sovereign exposure and other balance sheet characteristics are endogenous but *predetermined* variables. I will take this into account in the estimation by using the GMM Difference estimator. This dynamic panel data estimator employs a set of lagged internal instruments for endogenous, predetermined variables.

The coefficient in front of the sovereign shock in the first FOC is negative: it implies that as losses on sovereign bonds increase, equilibrium loan quantity decreases. The interest rate elasticity, α_0, which is assumed to be homogenous for all banks, enters into the coefficient of the sovereign shock. In the empirical strategy, I will somewhat relax the homogeneity assumption by interacting the shock with bank characteristics, such as size category (large vs. small) Aggregate factors $\lambda_{c,t}$ and r^f_t have natural proxies in country–time fixed–effects and η_b is a bank fixed–effect.

Note that, by substituting the solution for loan quantity into the loan demand schedule (1.6), a similar equilibrium condition for the loan interest rate ($r^L_{b,t}$) can be found. In this case, the sign of the coefficient in front of the sovereign shock ($\mathbb{X}(S_{b,t-1})$) is positive: an increase cost of funding for

[38] The only difference is that the regression will have $\Delta L_{b,t}$ rather than $L_{b,t}$ as dependent variable.

banks translate into an increase in the cost of capital for firms. This motivates the regression of the interest rate spread on syndicated loans.

To summarize, in this section I showed that, in a simple model of bank lending, tensions on the sovereign bond market, by increasing banks' cost of funding, decrease the loan quantity at equilibrium. According to this model, in order to identify a credit supply channel of banks' exposure to sovereign debt, loan demand needs to be homogenous across banks within the same country and time. Specifically, the demand shifter is at the country–time level or country–sector level in the empirical specification with loan interest rates. The loan interest rate elasticity is assumed to be the same for all banks or, at best, the same by size category (large vs. small).

Figure 1.1: GIIPS Sovereign Exposures, March 2010. EBA Stress Test 2010

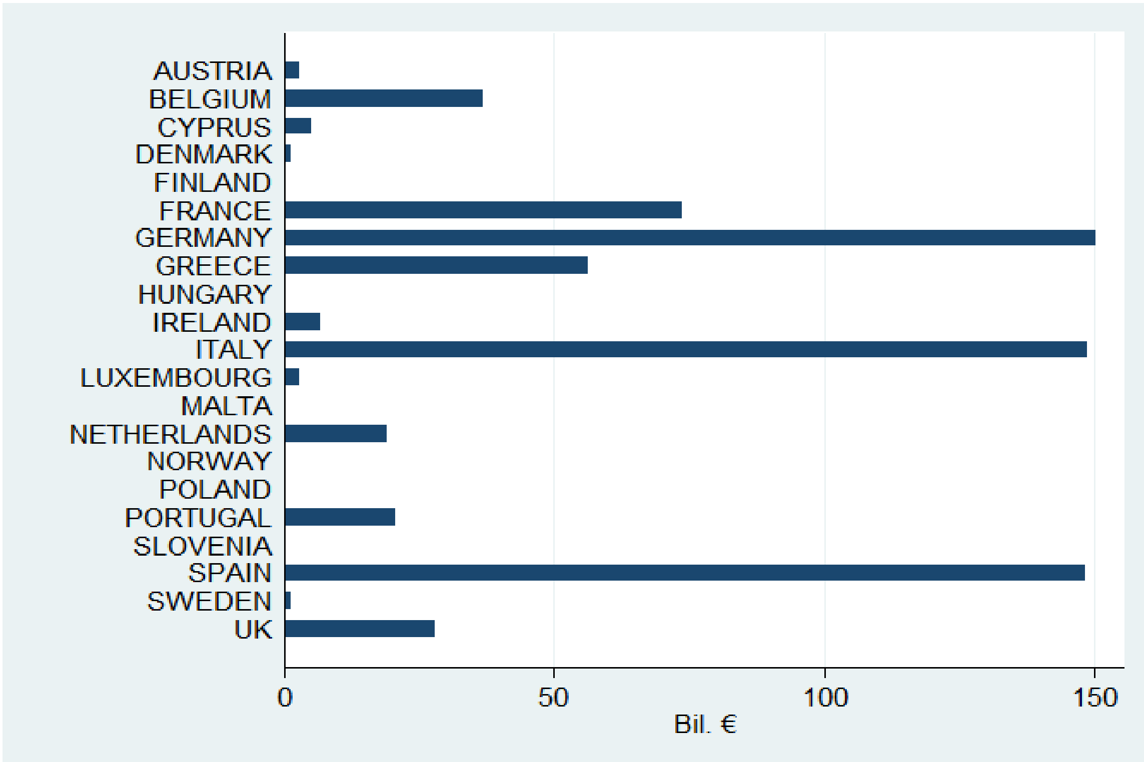


Figure 1.2: GIIPS Sovereign Exposures over Total Assets, March 2010. EBA Stress Test 2010

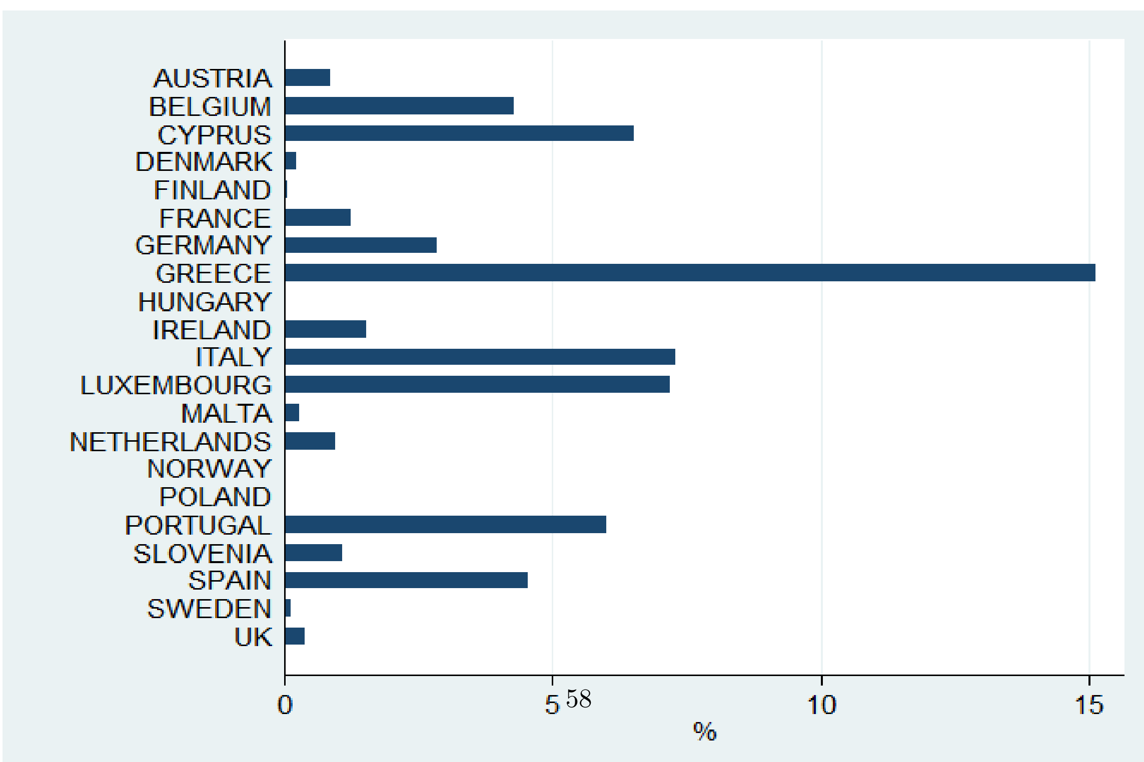


Figure 1.3: Domestic Loans Growth Rate to Non–Financial Corporations. ECB, MFI Aggregate Statistics

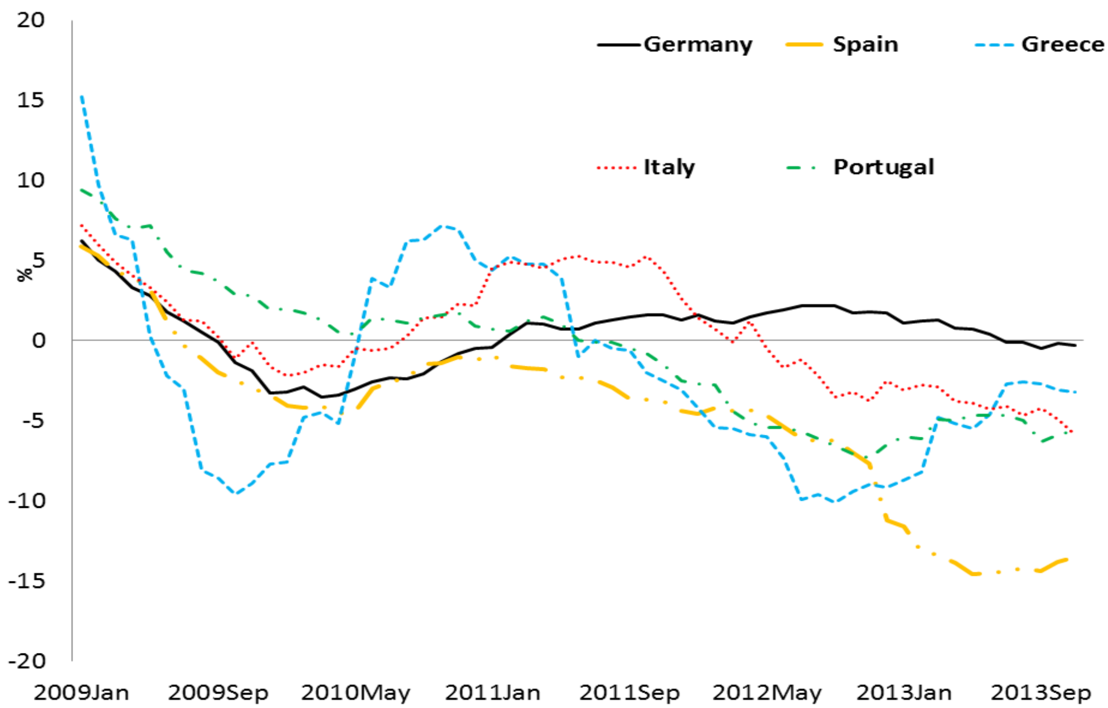


Figure 1.4: Average Interest Spread (on ECB Policy Rate) for New Loans to Non–Financial Corporations. ECB, MFI Aggregate Statistics (Narrowly Defined Effective Rates, all maturities and amounts).

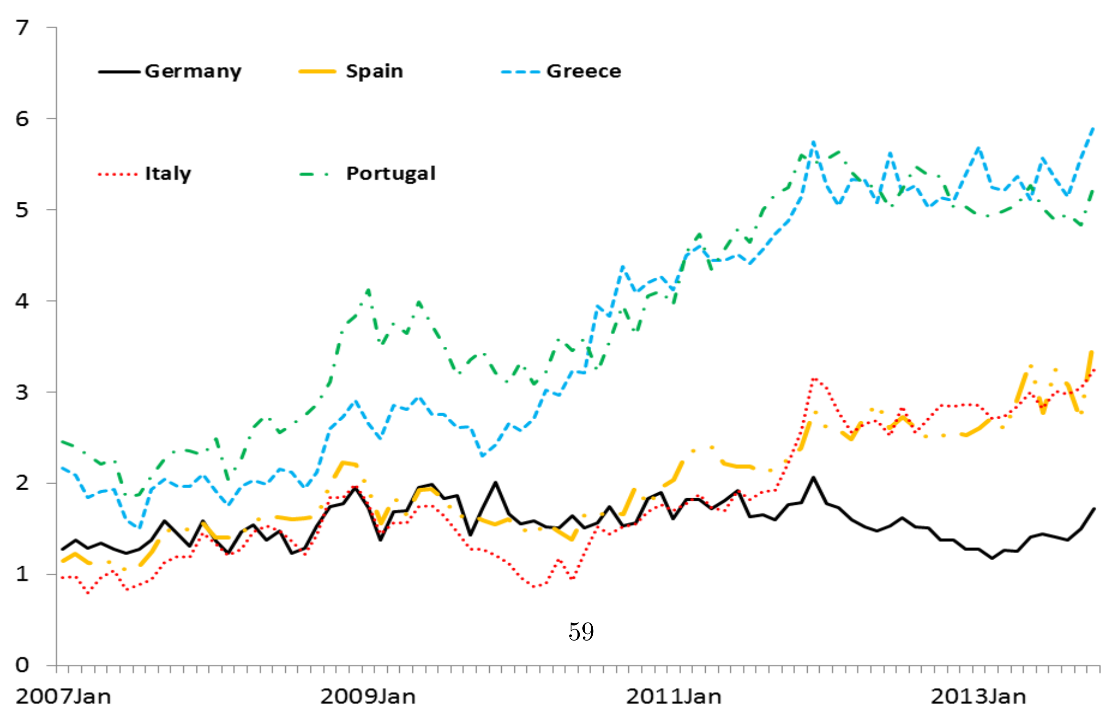
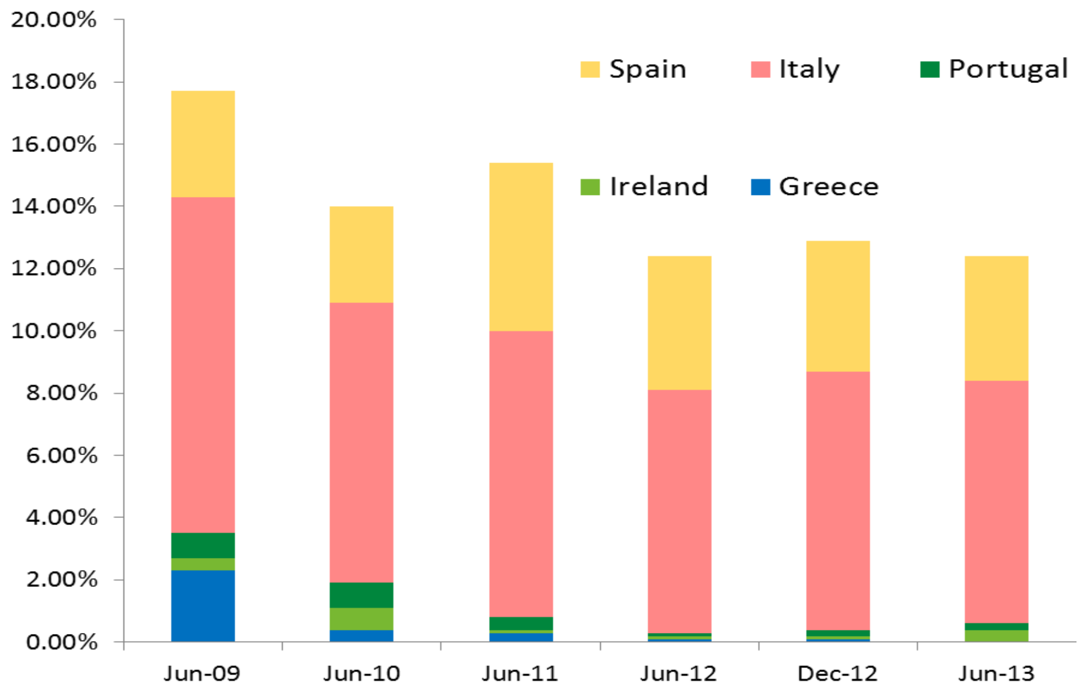


Figure 1.5: Share of GIIPS Collateral in Private GC Repo Market



Source: International Capital Market Association (ICMA) Survey, June 2013

Table 1.23: List of EBA banks

EBA CODE	**BANK NAME**	**CROSS BORDER**	**2009Q4 (2010Q1)**	**2010Q4**	**2011Q4**
AT001	Erste Bank Group (EBG)	Y	Y	Y	Y
AT002	Raiffeisen Bank International (RBI)	N	Y	Y	Y
AT003	Oesterreichische Volksbank	Y	N	Y	N
BE004	Dexia	Y	Y	Y	N
BE005	KBC Bank	Y	Y	Y	Y
CY006	Cyprus Popular Bank (Laiki)	Y	Y	Y	Y
CY007	Bank of Cyprus	N	Y	Y	Y
DE017	Deutsche Bank	Y	Y	Y	Y
DE018	Commerzbank	Y	Y	Y	Y
DE019	Landesbank Baden-Wurttemberg	N	Y	Y	Y
DE020	DZ Bank	Y	Y	Y	Y
DE021	Bayerische Landesbank	N	Y	Y	Y
DE022	Norddeutsche Landesbank	N	Y	Y	Y
DE023	Hypo Real Estate Holding AG	N	Y	Y	Y
DE024	WestLB	N	Y	Y	N
DE025	HSH Nordbank	N	Y	Y	Y
DE026	Helaba	N	Y	N	Y
DE027	Landesbank Berlin	N	Y	Y	Y
DE028	DekaBank	N	Y	Y	Y
DE029	WGZ Bank	N	Y	Y	Y
DE N/A	Deutsche Postbank	N	Y	N	N
DK008	Danske Bank	Y	Y	Y	Y
DK009	Jyske Bank	N	Y	Y	Y
DK010	Sydbank	N	Y	Y	Y
DK011	Nykredit	N	N	Y	Y
ES059	Banco Santander	Y	Y	Y	Y
ES060	BBVA	Y	Y	Y	Y
ES061	BFA-Bankia	N	Y	Y	N
ES062	La Caixa	N	Y	Y	Y
ES N/A	BASE	N	Y	N	N
ES083	CAM	N	N	Y	N
ES063	Effibank	N	N	Y	N
ES064	Banco Popular Espanol	Y	Y	Y	Y
ES065	Banco De Sabadell	N	Y	Y	N

ES066	DIADA - CatalunyaCaixa	N	Y	Y	N
ES067	BREOGAN - NovaCaixaGalicia	N	Y	Y	N
ES068	Mare Nostrum	N	Y	Y	N
ES069	BankInter	N	Y	Y	N
ES070	Espiga	N	Y	Y	N
ES071	Banca Civica	N	Y	Y	N
ES072	Ibercaja	N	Y	Y	N
ES073	Unicaja	N	Y	Y	N
ES074	Banco Pastor	N	Y	Y	N
ES N/A	Caja Sol	N	Y	N	N
ES075	Grupo BBK	N	Y	Y	N
ES076	UNNIM	N	Y	Y	N
ES077	Kutxa	N	Y	Y	N
ES078	Grupo Caja3	N	Y	Y	N
ES N/A	Caja de Cordoba	N	Y	N	N
ES079	Banca March	N	Y	Y	N
ES N/A	Banco Guipuzcoano	N	Y	N	N
ES080	Caja Vital	N	Y	Y	N
ES081	Caja de Ontinyent	N	Y	Y	N
ES082	Colonya	N	Y	Y	N
FI012	OP-Pohjola Group	N	Y	Y	Y
FR013	BNP Paribas	Y	Y	Y	Y
FR014	Credit Agricole	Y	Y	Y	Y
FR015	BPCE	N	Y	Y	Y
FR016	SocGen	Y	Y	Y	Y
GB088	RBS	Y	Y	Y	Y
GB089	HSBC	Y	Y	Y	Y
GB090	Barclays	Y	Y	Y	Y
GB091	Lloyds	N	Y	Y	Y
GR030	EFG Eurobank Ergasias	Y	Y	Y	N
GR031	National Bank of Greece	Y	Y	Y	N
GR032	Alpha Bank	Y	Y	Y	N
GR033	Piraeus Bank Group	Y	Y	Y	N
GR034	ATE Bank	N	Y	Y	N
GR035	Hellenic Postbank	N	Y	Y	N
HU036	OTP Bank.	Y	Y	Y	Y
HU N/A	FBH	N	Y	N	N
IE037	Allied Irish Banks	Y	Y	Y	Y
IE038	Bank if Ireland	Y	Y	Y	Y
IE039	Irish Life and Permanent	N	N	Y	Y

IT040	IntesaSanPaolo	Y	Y	Y	Y
IT041	Unicredit	Y	Y	Y	Y
IT042	Monte dei Paschi	N	Y	Y	Y
IT043	Banco Popolare	N	Y	Y	Y
IT044	Ubi Banca	N	Y	Y	Y
LU045	BCEE	N	Y	Y	Y
LU N/A	Banque Raiffeisen	N	Y	N	N
MT046	Bank of Valletta	N	Y	Y	Y
NL047	ING Bank	Y	Y	Y	Y
NL048	Rabobank	Y	Y	Y	Y
NL049	ABN AMRO	N	Y	Y	Y
NL050	SNS Bank	N	Y	Y	Y
NO051	DnB NOR Bank ASA	Y	N	Y	Y
PL052	PKO Bank	N	Y	Y	Y
PT053	Caixa Geral de Depositos	Y	Y	Y	Y
PT054	Millennium Bcp	Y	Y	Y	Y
PT055	ESFG	Y	Y	Y	Y
PT056	Banco BPI	Y	Y	Y	Y
SE084	Nordea Bank	Y	Y	Y	Y
SE085	SEB	Y	Y	Y	Y
SE086	Svenska Handelsbanken	N	Y	Y	Y
SE087	Swedbank	Y	Y	Y	Y
SI057	NLB Bank	Y	Y	Y	Y
SI058	NKBM	N	N	Y	Y
	TOTAL	**36**	**91**	**90**	**61**

Chapter 2

The Political Origin of Home Bias: the Case of Europe

2.1 Introduction

The European sovereign debt crisis has emphasized the importance of banks' exposure to sovereign debt. Banks' sovereign portfolios in Europe consist almost entirely of *domestic* government debt. The average (median) own exposure, defined as the proportion of domestic debt over the total sovereign portfolio, was 74% (86%) at the end of 2010. Figure 2.1 reveals that there is a significant degree of heterogeneity in banks' holdings of domestic debt within Europe: the median own exposure is in general higher in the periphery (PIIGS) than in Northern Europe, with Germany (DE) being a notable exception. The level of domestic exposures is well in excess of what standard finance theory would predict: there is significant *home bias* (Figure 2.2) [1]. In general, the home bias in sovereign bonds among European banks is quite persistent over time (Figure 2.3) and it tends to be higher

[1] The Capital Asset Pricing Model (CAPM) predicts that, in frictionless financial markets, homogenous investors would hold a share of financial asset equal to the share of the financial assets of that country in the world portfolio (see Cochrane (2005), page 155, and Coeurdacier and Rey (2012)). In the context of sovereign bonds, we use the

in the periphery than in Northern Europe [2]. The sovereign exposures have important implications for the real economy. During the sovereign debt crisis, banks incurred losses owing to the decline in market value of the sovereign debt in their balance sheets and some recent papers have shown that these potential losses are responsible for a sizable portion of the decline in lending.[3] Given that the riskiest bonds during the sovereign debt crisis were those issued by the PIIGS, the large degree of home bias is especially troubling among PIIGS banks and it may have exacerbated the recent recession.

In this paper, we investigate why European banks display a significant home bias in sovereign bond holdings. We believe that certain banks hold a disproportionate amount of their own country's sovereign debt because they are controlled by domestic politicians. In fact, many European banks have an explicit political participation through a block of shares owned by either the regional state in which the bank is headquartered (Germany) or the national government (Spain, Sweden, Portugal) or an indirect control exerted through private foundations whose directors are appointed by local or national politicians (Italy). These politicians may be interested in financing discretionary public spending to maximize their own objectives. These objectives may be increasing the chances of re–election or diverting public funds to friends, relatives or controlled firms. They would then persuade the politically controlled banks to finance national or local state borrowing by purchasing government bonds [4].

We define our "political influence" variable as the total percentage of shares held by central or

home bias measure as defined in Coeurdacier and Rey (2012):

$$HomeBias = 1 - \frac{\text{Share of Foreign Sovereigns in Bank } i \text{ Sovereign Holding}}{\text{Share of Foreign Sovereign Bonds in the Global Portfolio}}$$

When the home bias measure is equal to zero there is perfect diversification; when it is equal to one there is perfect home bias. Anything in excess of zero indicates some level of home bias.

[2] Ireland is a noticeable exception, with an extremely volatile home bias. We believe this may be due to the fact that these are raw data that do not take into account changing composition at the bank level (mergers & acquisitions and bank failures).

[3] See Bofondi et al.(2013), Popov and Van Horen (2014).

[4] The European Banking Authority (EBA) data we use unfortunately do not distinguish between central and local government debt. The sovereign exposure we refer to in this paper are the sum of the two.

local governments or by political foundations in the *pre–crisis* period (2006 or 2009). The reason for using pre–crisis data is twofold: first we want to avoid biases given by bank nationalizations that occurred in 2010–2011. In fact, during the crisis, many governments, especially in Germany, Spain and the UK, were forced to intervene to recapitalize or bail–out insolvent banks. Thus, if we were to measure "political influence" in 2011, we would largely overestimate the state's presence in the banking sector. Second, we claim that, although clearly endogenous, public ownership of these banks is a historical, predetermined presence, thus unlikely to be correlated with the error term in our main regression. The hypothesis we take to the data is that a bank that has a historically strong political presence among its shareholders will purchase more domestic sovereign bonds relative to a bank that does not receive any political pressure. We find evidence for this hypothesis both before and during the European sovereign debt crisis: a bank above the median political control has, ceteris paribus, a home bias of 10 to 19 percentage points higher than a bank below the median.

Moreover, we exploit the fact that there have been plenty of equity injections by each member State in the domestic banking system in both 2010 and 2011 to document the extent of *political pressure/moral suasion* on the controlled banks. We show that, during the sovereign debt crisis and upon receiving government support, banks significantly increased their exposure to domestic sovereign bonds *only if* they have strong political affiliations. The effect is twice as large for "political" banks located in the periphery (+8.9%) than for other "political" banks located elsewhere in Europe (+4.2%). These results are not explained by other country–time factors such as higher sovereign yields in the periphery, that would encourage purchase of sovereign bonds bonds from peripheral banks.

But government equity injections may be specifically targeted to political banks. If this were the case, it would not be surprising that only political banks buy domestic bonds after receiving government help. However, this does not appear to be the case: in 2010 and 2011, equity injections were not directly targeted to banks that have larger political affiliations, but rather to banks that performed worse in the previous year in terms of lower profitability and a larger pool of non performing loans. In other words, although governments seem to "twist arms" of politically controlled institutions into buying domestic debt, they also provide financial aid to those that actually need it.

The *political channel/moral suasion* hypothesis is not the only explanation for the home bias in sovereign bonds, both before and during the sovereign debt crisis. Standard information asymmetries arguments, where the local investors are better informed about the domestic sovereign than foreign investors, or other hedging motives may still be present. More recently, several papers explain the increase in home bias in the PIIGS countries during the sovereign debt crisis with creditor discrimination theories (Broner et al. (2014) and Brutti and Saure (2013) or as arbitrage opportunities fueled by the ECB LTROs (Acharya and Steffen (2014)). We do not challenge these hypotheses, we only show that politics also plays a role.

There is yet another reason for banks to hold sovereign bonds, which is sometimes refer to as *capital arbitrage*. Under Basel II, government bonds are considered almost risk free[5] so that banks would load their balance sheets with sovereigns to reduce risk–weighted assets and increase capital ratios. We argue that while the regulatory framework has certainly been a key factor in the excessive sovereign exposure among European banks, it cannot explain home bias. In fact, the zero risk weight applies not only to domestic sovereign debt, but to all countries in the European Union.

The paper proceeds as follows. Section II reviews the related literature. In Section III, we highlight some country-specific institutional details that are relevant to our analysis; Section IV describes the data and the methodology used in the paper. Section V presents the results and Section VI concludes.

[5] According to Basel II regulation, in order to compute Risk Weight Assets (RWA), banks can use two approaches: the Standardized Approach and the Internal–Ratings Based (IRB) approach. According to the first, government bonds receive a 0% risk weight as in Basel I. Under the IRB instead, the weight should be strictly positive, because, even though the model may assign a very low probability of default (PD) to a sovereign issuer, the loss given default (LGD) is positive. In practice, PD on sovereign debt are equal 0.1% for 201 major international banks (BIS Quarterly Review, December 2013). Moreover, in the European Union, there is a loophole that allows banks using the IRB to switch back to the Standardized Approach when evaluating sovereign bonds ("IRB permanent partial use").

2.2 Literature Review

The relationship between home bias in sovereign holdings and political influence at the bank level, is, to the best of our knowledge, not been established in the literature.

First, we contribute to the enormous international finance literature pioneered by French and Poterba (1991) that studies home bias in portfolio holdings. Most papers in this area have documented home bias in equity rather than bond holdings. A few exceptions are Bertaut et al. (2013), who show the decline in financial bonds' home bias among U.S. investors and Lane (2006), who shows that member countries of the European Economic and Monetary Union disproportionately invest in one another and especially towards their trade partners. Several recent papers analyze the increase in sovereign home bias among banks during the recent sovereign debt crisis. Battistini et al. (2014) document that only PIIGS banks respond to increases in country risk by increasing their exposure to domestic sovereign bonds, while banks from core countries do not, suggesting that *redenomination/repatriation* risk, *i.e.* the risk that the liabilities of banks would be renominated in the local currency, is the driving force behind the increase in home bias. Becker and Ivashina (2014) document a positive correlation at the country level between domestic government holdings by national banks and aggregate measures of state ownership in the banking system. Brutti and Saure (2013) analyze cross–country evidence in favor of the secondary market theory suggested by Broner et al. (2014). According to this hypothesis, in a crisis period, domestic banks would buy domestic debt in the expectation that the government will not default on domestic creditors. All these papers analyze country level data, which ignores the cross–sectional heterogeneity at the bank level. Using a bank-level panel, and constructing a measure of direct government ownership at the bank level, we are able to dig deeper and provide an explanation for why some banking institutions hold a disproportionately large amount of domestic sovereign bonds over total sovereigns.

Some theoretical papers also explore the home bias issue. Diamon and Rajan (2011) advance the hypothesis that banks are keen to load up with illiquid assets because the states of nature in which these assets default is the same in which the bank itself goes bankrupt; in other words, banks rationally put all their risk in a state of the world that would be catastrophic for them anyways (*risk synchronization*). Their argument has a natural application if one considers sovereign bonds an illiquid asset. Acharya and Rajan (2013) and Crosignani (2014) show that myopic governments have incentives to increase risk synchronization. The evidence we find suggests that, upon receiving liquidity injections, only the "political banks" boost their exposure to domestic government bonds relative to foreign ones, thus synchronizing even more their default risk with that of their respective domestic country.

Our findings also contribute to the literature on the performance of state-owned banks and to the literature on related lending. Barth (2001) provide a broad overview on the effects of regulation and ownership structure on the performance of the banking system. In general, they find that greater state ownership of banks tends to be associated with less developed banks and financial markets. Sapienza (2004) finds that firms located in areas where the party of affiliation of the bank's chairman is stronger receive more favorable loan conditions; Cuñat and Garicano (2010) show that banks whose chairman held a political position in the past perform worse than other banks. La Porta et al. (2003) find evidence that loans extended to related parties, either family members or controlled firms, have on average lower rates, lower collateral requirements and are more likely to default than unrelated ones. Finally, Khwaja and Mian (2005) document that in Pakistan government banks extend a preferential treatment to politically connected firms.

2.3 Institutional Details

There are three European countries that stand out for the pervasive and systematic role of politicians and local governments in the management of banks: Italy, Germany and Spain. In what follows we provide some key features that distinguish each of these countries in terms of political presence in the banking system. We also discuss the case of France as an example of a banking sector without any direct political influence, at least in the last decade.

Italy

Following the wave of liberalizations and privatizations that started at the European level in the 1980s, the Bank of Italy and the government made an attempt to privatize the numerous state-owned banks. In 1990, the Amato–Carli law transformed the state-owned banks into private entities; these were controlled by Foundations (non–profit organizations), that would have had to place their shares on the market at a later date to complete the privatization. In 1998, however, the Amato–Ciampi law superseded the 1992 law, reiterating that Foundations are non-profit organizations, but adding that these should operate under private law and not under public law as in the previous regime. As they became private entities, they could no longer be forced to progressively sell off their shares on the market. Thus, Foundations were able to maintain their controlling stakes in most Italian banks to the present day.[6] Importantly for our purposes, even though they became private, non-profit entities, they are still under the influence of political groups. The members of the board of directors in the Foundations are often appointed by local or national politicians.

Apart from one specific bank, MPS, the other four Italian banks in our dataset present more than one Foundation among their shareholders. Moreover, most of the times, within each Foundation there are members coming from both left and right wing parties. This degree of heterogeneity should convey the idea that, in the majority of the cases, banks with a large concentration of Foundations are influenced by a wide range of political parties.

[6] Boeri (2012) http://www.lavoce.info/i-politici-ai-vertici-delle-fondazioni-bancarie/ (in Italian)

Germany

The German banking system is organized in three different "pillars": private banks, such as Deutsche Bank and Commerzbank; cooperative banks, based on a member-structure where each member has one vote and, finally, *public banks*.

The latter are financial institutions, typically owned by the regional states (*Lander*) or by administrative districts or cities in which they are headquartered. Among these there are savings banks (*Sparkassen*) whose shareholders are usually local municipalities and the regional banks (*Landesbanken*), that are mostly owned by their respective *Lander* through a regional savings bank association.

Thus, in the case of Germany the definition of "political banks" is clear: those that have a direct state participation among their shareholders. In the EBA dataset, the political banks will mostly be the *Landesbanken*.

Spain

Savings banks represent a fundamental pillar of the Spanish banking system: founded in the 18*th* century with the objective of channeling private savings towards socially beneficially investments, savings banks accounted for 40 percent of Spanish banks' total assets in 2010.

They became financial institutions that do not distribute profits, that have no formal owner, but several governing bodies representing two different classes of stakeholders: *insiders* and *outsiders*. Insiders are employees, depositors and private founders; outsiders are the regional governments and other public entities. The relative voting power of the two groups in each bank depends on the specific regional law. Around a decade ago, in 2002, a national reform capped the representation of public entities, including regional governments, at *50 percent of the voting rights* in each bank [7].

[7] In July 2010, the ceiling on voting rights of public entities was reduced to 40 percent and professional expertise was required to sit in a governing bodies.

Since outsiders, *i.e.* the public entities, are focused on achieving socially oriented goals, improving profitability has not always been the main objective of savings banks. In this regard, **?** document that Savings banks were more likely to open new branches and extend new loans in provinces that were politically close; additionally, **?** provide some evidence that savings banks whose chairman has political affiliations performed worse than other banks.

France

A different path has been followed by France.[8] After World War II and up to the late 1980s, almost all banks, both investment and commercial, were either state-owned or co-operatives. The Chirac government changed the situation when, in 1987, he privatized several major banks, including *Societe Generale* and *Paribas*. Another wave followed few years later in 1993, with the privatization of BNP among others. The complete privatization of the banking system was accomplished in 2001.

2.3.1 Theory of Inter-Party Support

Whereas the political influence in the case of direct state ownership (Germany, Spain) is clear, the case of foundations' ownership (Italy) requires a more careful analysis: certain banks are affiliated with only one political party that is not necessarily in power at any given point in time. Monte Dei Paschi (MPS) in Italy is an example of such a bank: it is affiliated with the centre–left municipal government and it has a strong home bias (96%) even in 2010, when the national government is from centre–right. For these institutions, it is not clear why a political party that is not ruling the country would be interested in buying sovereign bonds and finance public spending of their opponents. It may be interested in doing quite the opposite in order to destabilize the incumbent government.

We claim that there are two main reasons that may explain that behavior. First of all, local

[8]Alain Plessis (2003), The history of banks in France, *Federation Bancaire Francaise*.

politicians can sustain a central government of the opposite political affiliation in exchange for monetary transfers to the respective region or local municipality. There is suggestive evidence, for the case of Italy, that regional transfers are not primarily dictated by shared political affiliation; a more crucial determinant is the political strength of the party in the specific region, regardless of political affiliation.[9]

Second, a theory that supports inter-party funding is borrowed from the political science literature. Katz and Mair (1995, 2009) are the first to document that political parties in a wide range of developed countries have started to behave like a cartel. Instead of competing against each other on relevant issues, they transfer more and more competences upward to technocratic and non-partisan commissions. Perhaps more importantly, they decided to alter the structure of payoffs: they agreed on the introduction of public financial subventions to political parties that are guaranteed regardless of whether a party wins or loses. This last piece of regulation severely limits the incentives to compete in order to win the elections, as the monetary payoffs are not linked to the election's outcome. Hence, the concern that a bank affiliated with a leftist party would have the incentive to destabilize the governing right wing party is clearly downsized in light of the findings of Katz and Mair.

2.4 Data and Methodology

Data

The dataset is the result of the merger of three different sources: detailed bank level data on the exposure to sovereign bonds and liquidity injections from the EU–wide Stress Test and Recapitalization Exercises; information on the *pre–crisis* degree of political presence for each bank is collected

[9] Greco (2009) (mimeo).

from Annual Reports whenever available; other balance sheet data comes from Bankscope.

The key dataset contains the information on political influence. Specifically, we want to create a variable that captures the degree of government or politicians' control within each bank. We also want this measure to be dated *prior* to the sovereign debt crisis so as not to bias the degree of political control with any bail out or nationalization policy that occurred in 2010–2011. Hence we collect 2006 Annual Reports for all banks, except Spanish ones. It is in fact difficult to find Annual Reports at earlier dates for some small Spanish Savings banks, because many were recently acquired or merged with other banks. For this reason, we resort to IMF data, which list for each savings bank the percentage of voting rights held by local governments prior to 2009.[10] We construct the variable *Political* as the sum of any participation held by the local or central governments and by political foundations in each bank; we then normalize it by its standard deviation. We sometimes find useful to use a dummy variable, $\mathbb{1}$*(Political)*, which takes value of one if the degree of political control in a specific bank is above the median of the domestic country and zero otherwise. We also create a dummy variable, $\mathbb{1}$*(Cooperative)*, which takes value of one if the bank is a cooperative and zero otherwise. It is important to distinguish between a non–cooperative and cooperative bank, since the latter display certain features that are similar to those of highly politicized banks even though they are not owned by the state (diffuse ownership among cooperative members). Notably, cooperative banks display a high degree of home bias in sovereign bonds, similar to political banks, probably due to the very "local" nature of their business model.

Sovereign exposure data have been collected by the EBA in the context of the EU–wide Stress Test and Recapitalization Exercises. Specifically in an effort to enhance transparency and restore confidence in the financial system, the EBA decided to disclose bank-by-bank result for both the 2010 and 2011 Stress Test Results and the so-called 2011 and 2012 Recapitalization Exercises. These

[10] ?

exercises contain information on the capital composition, including government aid in the form of equity support measures, credit risk exposure and sovereign debt exposure to each of the 30 members of the European Economic Area (EEA 30) for all the participating banks. The sample consists of 90 European banks in March 2010, which we will refer to as end–of–year 2009, and December 2010; 61 banks in December 2011 and June 2012, covering at least 60% of banking assets in Europe and at least 50% in each Member State. For December 2010 only, a breakdown of the credit portfolio by categories of borrowers is available. For instance, we know the amount of credit granted to private corporations, public institutions, small and medium enterprises, the exposure to residential mortgages and the amount of defaulted loans.

Finally, we match the EBA dataset with banks' balance sheet data obtained from Bankscope.

Empirical Methodology

To measure the effect of political presence in a bank on its degree of home bias in sovereign bonds, we run a set of *cross-sectional* regressions for 2009, 2010, 2011.[11] For each year we employ the following specification:

$$HomeBias_i = \beta_1 Political_i + \gamma' X_i + \mathbb{1}(Cooperative_i) + D_i + \varepsilon_i \tag{2.1}$$

where $HomeBias_i$ is one of two measures: i) the ratio of domestic bonds held by bank i over the total European debt ($\frac{Own}{TS}$) [12] or ii) the home bias measure in Coeurdacier and Rey (2012). The first measure is the most intuitive, but it ignores the nominal size of each country's debt. For example, it is reasonable for Italian and German banks to have a larger exposure to their home country's debt

[11] The main variable of interest $Political_{i,j}$ does not vary over time, so we cannot use panel regressions with bank fixed effects.

[12] The EBA sovereign exposure data contains only countries belonging to the European Economic Area (EEA30), a group of 30 countries which broadly coincides with the European Union. Only in December 2010 exposure to US and Japan was disclosed, but we drop these countries from our analysis as they are only available for one year.

than Belgian and Dutch banks because Italian and German public debt are much larger. However, this does not pose a problem in estimation because we control for country fixed effects that also absorb the size of a country's debt.The second measure, on the other hand, explicitly takes this into account. It is defined as follows in Coeurdacier and Rey (2012):

$$HomeBias = 1 - \frac{\text{Share of Foreign Sovereigns in Bank } b \text{ Sovereign Holding}}{\text{Share of Foreign Sovereigns in the Global Portfolio}}$$

where *Global* is represented by the EEA30 countries in our data. This measure is bounded between zero (perfect diversification) and one (perfect home bias), while anything in excess of zero indicates some level of home bias.[13] Comparing Fig. 2.1 and Fig. 2.2 reveals however that the difference between the two measures is negligible: neither the country ranking nor the level of home bias is very much affected. For example, Italian and German banks have high positive values in both cases.

$Political_i$ is also one of two measures. It is either the percentage of shares owned by the domestic government or other domestic political entities (Foundations in Italy, for example) divided by its standard deviation; or it is a dummy, $\mathbb{1}(Political_i)$, equal to one if the bank is above the median of the distribution of government ownership in each country.

Other explanatory variables we use are: X_i, a set of *lagged* bank balance sheet characteristics (log of total assets, Tier1 ratio, Leverage, Deposits over Total Funding, ROAA, Non Performing Loans over Gross Loans) and $\mathbb{1}(Cooperative_i)$ a dummy equal to one if bank i is a cooperative bank. We allow for a different intercept in home bias for cooperative banks because these banks are characterized by dispersed ownership among members (one–head–one–vote) and no share directly owned by the domestic government: nonetheless, they usually exhibit a large home bias given by the very "local" nature of their business model. It is also possible that others, more indirect forms

[13] Note that the first measure of home sovereign bonds over total sovereigns does not exactly replicate this: it is also equal to one in case of perfect home bias, but it is equal to zero if the bank does not own any domestic debt, not if the bank is perfectly diversified. This difference turns out not to matter in the regression analysis.

of political influence are at play in cooperative banks: it is often the case, in Italy for example, that cooperative firms have strong ties with political parties. Finally, D_i a country dummy where the bank is headquartered. Country dummies are important because they control for country specific factors; more specifically, they take into account i) institutional characteristics and ii) optimal portfolio considerations.

The first motivation pertains to countries' institutional heterogeneity; for example, we need to control for the fact that in Spain the government participation in each bank, by law, cannot exceed 50%, whereas in Germany the local government can hold any number of shares. German *Landers* very often holds around 85% of shares in the *Landesbanken*. Since we are interested in evaluating whether a certain political ownership is large or small in a given country, a set of country–specific intercepts in the above regression is appropriate.

The second consideration has to do with asset pricing theory. The CAPM implies the following pricing equation: $1 = E_t\left[M_{j,t+1}R_{i,t+1}\right] = Cov_t\left[M_{j,t+1}, R_{i,t+1}\right] + E_t\left[M_{j,t+1}\right]E_t\left[R_{i,t+1}\right]$. $M_{j,t+1}$ is the stochastic discount factor or pricing kernel of country j at time $t+1$ and $R_{i,t+1}$ is the real rate of return on asset i at time $t+1$. The above equilibrium condition implies that the optimal holding of any asset, sovereign bonds included, depends on the covariance between a country specific factor and an asset specific component. Therefore, each bank in country j should have the same exposure to the set of sovereign bonds. For this reason, the set of country dummies also reflects country specific portfolio aspects. We use country dummies with both dependent variables, the home exposure over the total and the Coeurdacier and Rey (2012) home bias measure.

In our second specification, we exploit the time dimension of the panel to investigate how home bias varied *during* the crisis. We want to test the hypothesis that, upon receiving an equity injection from the domestic government, only political banks increase their home bias relative to other banks, especially if they are located in the periphery. This would be consistent with a *political* *pressure/moral suasion* hypothesis, where the domestic government calls on banks to buy sovereign debt at a time of low demand. Therefore, we run the following *panel* regression:

$$\Delta HomeBias_{i,t} = \beta_1 GovHelp_{i,t-1} + \beta_2 \big(\mathbb{1}(PIIGS) \times \mathbb{1}(Political_i)\big) GovHelp_{i,t-1} + \\ + \beta_3 \big(\mathbb{1}(NOPIIGS) \times \mathbb{1}(Political_i)\big) GovHelp_{i,t-1} + \gamma' X_{i,t-1} + \eta_i + \lambda_t + \varepsilon_{i,t} \tag{2.2}$$

where $\Delta HomeBias_{i,t}$ represents the change in home bias of bank i at time t, $GovHelp_{i,t-1}$ is the amount of equity injection given by the domestic government to bank i at the beginning of the year as a fraction of Risk Weighted Assets (RWA). $\mathbb{1}(Political_i)$ is a non–time varying dummy if bank i is above the median political control in each country and $\mathbb{1}(PIIGS)$ $\mathbb{1}(NOPIIGS)$ are dummies for whether bank i belongs to PIIGS or not. In this regression we allow for the effect of equity injections to differ depending on whether a bank is politically influenced and *at the same time* whether or not it belongs to the PIIGS. Finally we control for bank fixed–effects, η_i, and for either year– or country–year fixed–effects λ_t.

Table 2.1 below reports some summary statistics of the dataset.

Table 2.1: Summary Statistics at December 2010

	N	Mean	St.Dev.	Min	Max
Political ownership in 2006, (%)	90	21.3	31.3	0	100
$\mathbb{1}(Political_i)$ in 2006	90	0.34	0.47	0	1
Home Bias (Own/TotSov), (%)	90	74	26.4	9.8	100
Home Bias (CR(2012)), (%)	87	71.7	29	.05	100
Gov Help/RWA, (%)	90	1.4	3	0	21.85
$\mathbb{1}(GovHelp > 0)$	90	.38	.46	0	1
Gov Help/RWA if > 0, (%)	35	3.6	4	.32	21.84
Tier1/RWA, (%)	90	11	3.7	4.3	34.7

On average, the pre–crisis (2006) ownership by domestic government or political entities in Europe is at 20% among the 90 banks participating in the European Stress Test in 2010. However, only 48 banks (53% of the sample) have at least some level of political ownership and the dummy $\mathbb{1}(Political_i)$ shows that only 34% of the sample can be classified with an above the median political control in each country. Home bias is high on average according to both measures, however there is also a large heterogeneity, as was evident from Figures 2.1 and 2.2. Also, 38% of the banks in our sample received some form government help at the end of 2010, with an average, conditional on the help being positive, of 3.6% of RWA. These are big numbers considering that, on average, the Tier1 over RWA ratio is at 11% for all banks.

2.5 Results

Table 2.2 above reports the results for the main set of cross-sectional regressions. In the first three columns we regress the exposure to domestic sovereigns over total sovereigns in 2009, 2010 and 2011 on our continuous variable for political influence, *Political*, and a set of controls. In the last three columns we repeat the exercise but now we use a dummy variable, $\mathbb{1}$(Political), to capture political influence within each bank.

From the first two rows we notice that the coefficients of interest are always positive and significant at 5%.[14] This implies that banks that are more politically influenced display greater home bias in sovereign bond holdings. If we take a look at the first column, which we can think of as a pre-sovereign debt crisis regression, we have that a one standard deviation increase in the level of political influence (30 pct.points) is associated with an increase in the domestic composition of the sovereign bond holdings by 8%, which is about one third of a standard deviation of $\frac{Own}{TS}$.

[14] Only in one out of six cases, the coefficient on $\mathbb{1}(Political_i)$ is significant at 10%. Notice how it is significant at 1% using the $Political_i/std.dev$ measure in 2009 too.

Table 2.2: Home Bias (Own/TotalSovereign) and Political Presence.

	HB_{2009} (1)	HB_{2010} (2)	HB_{2011} (3)	HB_{2009} (4)	HB_{2010} (5)	HB_{2011} (6)
Political$_i$/std.dev.	7.747***	5.867***	5.717***			
	(3.34)	(2.71)	(3.28)			
$\mathbb{1}$(Political)$_i$				11.86*	10.19**	15.89***
				(1.95)	(2.41)	(4.11)
$\mathbb{1}$(Cooperative)	18.11***	16.38**	10.16	18.51**	16.74**	12.42**
	(2.92)	(2.67)	(1.68)	(2.54)	(2.62)	(2.20)
log(Asset)$_{t-1}$	-7.241**	-10.13***	-14.47***	-7.779***	-10.24***	-14.99***
	(-2.47)	(-3.94)	(-4.38)	(-3.95)	(-4.25)	(-5.40)
Tier1$_{t-1}$	1.426	-2.520**	-0.902	1.014	-2.256*	-0.973*
	(0.88)	(-2.12)	(-1.63)	(0.64)	(-1.85)	(-1.89)
Leverage$_{t-1}$	-0.350***	0.0892	-0.248	-0.398***	0.0495	-0.224
	(-3.26)	(0.69)	(-1.61)	(-3.51)	(0.39)	(-1.55)
(Dep/TF)$_{t-1}$	-0.513**	0.134	0.0195	-0.556**	0.141	0.0213
	(-2.14)	(0.62)	(0.05)	(-2.06)	(0.60)	(0.05)
ROAA$_{t-1}$	-0.113	5.995	-0.0235	-0.0733	4.018	-1.633
	(-0.03)	(1.18)	(-0.01)	(-0.02)	(0.84)	(-0.52)
(NPL/GL)$_{t-1}$	3.540**	-0.816	-1.139	4.091**	-0.613	-1.054
	(2.35)	(-0.68)	(-1.64)	(2.34)	(-0.54)	(-1.66)
N	71	77	57	71	77	57
Country Dummies	yes	yes	yes	yes	yes	yes

t statistics in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note: set of cross–sectional regressions of banks in 2009, 2010, 2011. The dependent variable is the ratio of domestic sovereign over total sovereign. $Political_i/std.dev$ is the ratio of political ownership over its standard deviation; $\mathbb{1}(Political)_i$ is a dummy equal to one if the bank is above the median of the distribution of political ownership in each country; $\mathbb{1}(Cooperative_i)$ is a dummy equal to one if bank i is a cooperative; log($Assets_{t-1}$), $Tier1_{t-1}$, $Leverage_{t-1}$, $(Dep/TF)_{t-1}$, $ROAA_{t-1}$, $(NPL/GL)_{t-1}$ are, respectively, the log of total assets, the Tier1 ratio over RWA, the ratio of total assets and common equity, the deposits to total funding ratio, the average return on assets and the ratio of non–performing loans over total loans, all lagged by one year. Std.err. are White HAC robust.

On the other hand, by looking at the fourth column, the coefficient on the political dummy implies that a bank that moves from the bottom 50% to the top 50% of the distribution of political influence displays on average 12% more weight to domestic sovereigns relative to the total, which is about half of a standard deviation of $\frac{Own}{TS}$.

A covariate that is always highly significant is bank size, as measured by the log of total assets. The coefficient is a semi–elasticity and it implies that for a 1% increase in total assets, the own exposure is expected to decline by 0.1 percentage points on average across all the years. The punchline is that larger banks have a smaller home bias in sovereign bonds: the sovereign portfolios of larger institutions are more diversified. Also cooperative banks, on average and all else equal, have a own exposure 13 to 17 percentage points higher than other banks, at least before 2011. Note that cooperative banks have no direct political or state ownership, but the significant degree of home bias may be explained by the very "local" nature of their business model. Or, possibly, it indicates that cooperative banks may be subject to other forms of indirect political influence, that our measure of political control cannot capture. It is often the case, at least in Italy, that cooperative savings bank have strong political ties to political parties.

Table 2.3: Home Bias (Coeurdacier and Rey (2012)) and Political Presence

	HB_{2009} (1)	HB_{2010} (2)	HB_{2011} (3)	HB_{2009} (4)	HB_{2010} (5)	HB_{2011} (6)
$Political_i$	10.09^{***}	7.527^{***}	7.220^{***}			
	(3.81)	(3.01)	(4.05)			
$\mathbb{1}(Political)_i$				15.71^{**}	12.45^{**}	18.81^{***}
				(2.27)	(2.67)	(4.59)
N	69	75	58	69	75	58
Country Dummies	yes	yes	yes	yes	yes	yes
Other: Log(TA)(-)***, Coop(+)***, Tier1(-), Lev(-), NPL(-), Dep(-) and ROAA(+)						

t statistics in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Note: set of cross–sectional regressions of banks in 2009, 2010, 2011. The dependent variable is the **?** measure:

$HomeBias = 1 - \frac{\text{Share of Foreign Sovereigns in Bank } b \text{ Sovereign Holding}}{\text{Share of Foreign Sovereign Bonds in the Global Portfolio}}$

Other bank controls defined as before.

Next, in Table 2.3, we run the same regression but changing the dependent variable to the **?** measure. The main results are basically unchanged, if anything both the magnitude and the significance of the coefficients is larger in this specification.

2.5.1 Panel Regression: Political Pressures during the Sovereign Debt Crisis

Now we ask whether politicians exerted pressures on controlled banks during the sovereign debt crisis; more specifically, we want to test whether, upon receiving equity injections, political banks increased their exposure to domestic sovereign bonds. We expect the effect to be stronger for banks in the PIIGS, where the respective governments had an incentive to encourage the purchase of government bonds so as to lower the yields.

Table 2.4 below summarizes this heterogeneous impact of liquidity injections of banks' portfolio decisions. It displays the panel regression of changes in exposure to domestic relative to total sovereigns. Government help here is defined as any form of equity injection, measured as a fraction of Risk Weighted Assets (RWA), given by the respective governments. The data come from the EBA Stress Tests and Recapitalization exercises, where either purchase of ordinary bank shares by the government or government support measures count as government help in the calculation of common equity of the bank. Column (1) suggests that receiving government help by itself does not affect the bank's choice between buying domestic or foreign sovereigns. However, a political bank that receives liquidity injections by the local government would increase its exposure to domestic sovereigns. The effect is larger if the bank is located in particularly distressed countries, namely the PIIGS. Column (2) indicates that an additional equity injection of 1% of risk weighted assets is associated with an increase in domestic relative to total sovereign exposure by almost 9%, compared to 4% for non PIIGS banks. The two coefficients are sufficiently precisely estimated so that a simple hypothesis test rejects the null that the two effects are the same at the 5% level. The results would look very similar if we used the alternative definition of home bias in Coeurdacier and Rey (2012).

These results suggest that sovereign countries, especially the PIIGS, use domestic political banks to purchase the bonds they issue when there is a lack of demand. Indeed, upon receiving freshly injected equity, only the politically controlled banks increase their degree of home bias.

Table 2.4: Political Pressure on the Banks. Panel regression.

	(1) $\Delta HomeBias_{i,t}$	(2) $\Delta HomeBias_{i,t}$
Gov Help$_{t-1}$	-.746	-0.739
	(-1.13)	(-1.12)
$\mathbb{1}(Political) \times GovHelp_{i,t-1}$	7.877***	
	(5.51)	
$\mathbb{1}(Political, Piigs) \times GovHelp_{i,t-1}$		8.981***
		(8.15)
$\mathbb{1}(Political, NoPiigs) \times GovHelp_{i,t-1}$		4.192**
		(2.08)
$N \times T$	187	187
N of banks	77	77
Bank + Year FE	yes	yes
Other Bank Controls: Tier1(-)***,Log(TA)(-)*,Lev(-),NPL(-),Dep(-)		
P-Value of the Test		
$\mathbb{1}$(Pol,Piigs)Gov = $\mathbb{1}$(Pol,NoPiigs)Gov		0.0337

t statistics in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Panel regressions. The dependent variable is $\Delta HomeBias_{i,t}$ defined as the change in the ratio of domestic sovereign bonds over total sovereigns between 2010Q4–2010Q1, 2011Q4–2010Q4, 2012Q2–2011Q4. $GovHelp_{i,t-1}$ is the government equity injection as a percentage of RWA given to bank i at the beginning of the period. Other variables are defined as before. Std.err. are clustered at the bank–year level.

The panel regressions in Table 2.4 are not controlling for the fact that, during the crisis, we may observe an increase in home bias because of country and time specific factors. In particular, it is conceivable that PIIGS banks may have decided to increase their home bias because of the very high yields in PIIGS sovereign bonds. Investing in these bonds was risky, but for PIIGS banks it may be perfectly rational to put all risk in a state of the world, a sovereign default, that corresponds to banks' defaulting themselves (*risk synchronization*). Also, these risky behaviors may have been funded by the ECB 3 year Long Term Refinancing Operations (LTRO) in December 2011 and February 2012 that injected large amounts of liquidity, borrowed at 100 and 75 bps. respectively, into participating banks (part of the *carry trade* hypothesis advanced by **?**). We test these hypotheses and the robustness of our results in Table 2.5 below.

Table 2.5: Political Pressure on the Banks: Robustness

	Yields and CDS (1)	Country–time FE (2)	LTRO (3)	Yields and CDS (4)	Country–time FE (5)	LTRO (6)
$\Delta Yield$	220.3			295.8		
	(0.77)			(0.98)		
$\Delta CDS5Y$	-254.9			-408.9		
	(-1.15)			(-1.49)		
LTRO/TotalAssets			-2.877			-3.360
			(-0.81)			(-0.93)
Gov Help$_{i,t-1}$	-1.149	0.369	0.489	-0.719	0.685	0.179
	(-1.62)	(0.48)	(0.66)	(-0.89)	(0.87)	(0.30)
$\mathbb{1}$(Political)×	5.050*	10.94**	12.17**			
Gov Help$_{i,t-1}$	(1.85)	(2.05)	(2.47)			
$\mathbb{1}$(Political, Piigs)×				7.343***	16.53**	16.53**
Gov Help$_{t-1}$				(2.26)	(2.42)	(2.08)
$\mathbb{1}$(Political, NoPiigs)×				4.096	4.814	11.15*
Gov Help$_{i,t-1}$				(0.72)	(1.36)	(1.67)
N	155	187	176	155	187	176
N bank	66	77	76	66	77	76
Bank + Year FE	yes	yes	yes	yes	yes	yes
Country–time FE	no	yes	yes	no	yes	yes

t statistics in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Panel regressions. The dependent variable is $\Delta HomeBias_{i,t}$ defined as the change in the ratio of domestic sovereign bonds over total sovereigns between 2010Q4–2010Q1, 2011Q4–2010Q4, 2012Q2–2011Q4. $GovHelp_{i,t-1}$ is the government equity injection as a percentage of RWA given to bank i at the beginning of the period. $\Delta Yield$, $\Delta CDS5Y$ are the growth rates of sovereign yields for 10 year bonds and 5 years CDS rates over the relevant periods. $LTRO/TotalAssets$ is the borrowing from the 3–year LTRO operation in December 2011 and February 2012 at the bank level (47 banks) over total assets. It is equal to zero for *all* banks before 2012Q2 and equal to the LTRO amount for the 47 banks for which information on the borrowed amount was found and missing otherwise. Other variables are defined as before. Std.err. are clustered at the bank–year level.

Columns (1)–(3) and (4)–(6) have the same set of controls, but the latter three split the effect of the interaction between government help and political banks among PIIGS and non–PIIGS banks. In particular, column (1) and (4) control for the *change* in the sovereign yield and CDS in each period.[15] We would expect the increase in the yield to increase home bias, but the increase in CDS, a proxy for risk of default, should offset it. The estimated coefficients are in fact positive and negative, respectively, but they turn out to be non–significant. Column (2)–(3) and (5)–(6) use a set of country–time FE that absorb all unobserved heterogeneity that is country and time specific, including yields and CDS. Finally, column (3) and (6) include bank specific usage of LTRO funds divided by total assets. Bank by bank figures on LTRO usage have not been released by the ECB, however we have collected data from banks' annual reports and industry reports for 47 major EBA banks. These banks borrowed €514 bn. in both LTROs, around half of total gross funds.[16] Figure 2.4 reveals that the LTRO have been dominated by Italian and Spanish banks (50% of the LTRO 1+2 funds), although, admittedly, the disclosure for French and German banks has been poor. In some cases we had to rely on industry estimates by Morgan Stanley Research (2012), because although it is known that a bank has participated, the actual amount was not disclosed on annual reports. Since both 3 year LTRO operations took place in December 2011 and February 2012, the LTRO variable takes a value of zero before 2012Q2 for *all* banks and then it is equal to the amount borrowed only for the 47 banks for which information is available (it is missing in 2012Q2 for the other banks).

The results in all columns show that our hypothesis is robust even controlling for country–time specific trends and the LTRO interventions: the coefficient on the interaction term between government help and political banks is significant in all specifications. Moreover, it appears that, after we take into account country–time characteristics, the political banks in the PIIGS are the only ones that increase home bias after receiving government help. It makes sense that the significance of the coefficient survives only for PIIGS banks, because these are the countries whose governments have

[15] The periods are: 2010Q4–2010Q1, 2011Q4–2010Q4, 2012Q2–2011Q4. The number of observations is differentin column (1) and (3) because Bloomberg does not provide data for the sovereign yields and CDS in all countries. We do not have information on Cyprus, Denmark, Finland, Hungary, Luxembourg, Malta, Poland and Slovenia.

[16] According to industry reports by Morgan Stanley Research (2012), only around half of gross funds were actually new *net funding*, as banks rolled over existing ECB facilities into the LTRO. The data we have collected are *mostly* on gross funds usage.

a higher incentive to pressure banks into buying domestic government bonds during the crisis.

2.5.2 Determinants of Government Help: Not just for Political Banks

One could think that our previous set of results, the fact that political banks buy more domestic government bonds during the crisis, could be explained by the fact that only political banks received government help during the crisis. In that case it would not be surprising that only political banks increase their respective own exposure to domestic governments. In Table 2.6 we show that this is not the case. We estimate the relationship between the amount of equity provided by local governments to each bank and a set of regressors, including past performance and political influence. The punchline is that equity injections are targeting banks with low prior profitability and larger pools of non-performing loans, not political banks directly. However, upon receiving these injections, only politically controlled banks increase their holdings of domestic sovereigns.

The first two columns report OLS regressions while the last two use the Tobit estimator (we report the slope coefficients in both cases). The last approach is more appropriate for this scenario because we should think of government help as a censored variable. It is equal to zero if the bank is in good shape and the government decides not to extend support or any positive value otherwise. The right specification that takes in to account both the discrete choice of whether or not to support a bank and the magnitude of the liquidity injection is the Tobit model.

The qualitative outcome of the OLS and the Tobit regressions are the same: political influence has not played any additional role in attracting more support from the government. The two main factors associated with greater government help are lower profitability (Return on Average Assets, ROAA) and more non-performing loans over gross loans (NPL/GL).

By looking at the marginal effects (not reported in the table) for 2010, we see that a decrease in

Table 2.6: Determinants of Government Help

	(1) OLS Gov Help$_{2010}$	(2) OLS Gov Help$_{2011}$	(3) Tobit Gov Help$_{2010}$	(4) Tobit Gov Help$_{2011}$
$\mathbb{1}$(Political)$_i$	-0.572	-0.560	-0.462	-0.000265
	(-0.86)	(-0.62)	(-0.37)	(-0.00)
ROAA$_{t-1}$	-1.880***	-1.361***	-5.222***	-1.928***
	(-3.16)	(-6.81)	(-4.96)	(-3.95)
(NPL/GL)$_{t-1}$	0.300**	0.146**	0.935***	0.525***
	(2.60)	(2.01)	(2.90)	(2.84)
$\mathbb{1}$(Cooperative)	-2.360***	-1.100	-6.226**	-2.210
	(-2.71)	(-1.58)	(-2.46)	(-0.88)
$\mathbb{1}$(PIIGS)	0.172	1.488	-0.0145	0.653
	(0.37)	(1.56)	(-0.01)	(0.30)
log(Asset)$_{t-1}$	0.114	-0.160	0.753*	-0.316
	(0.56)	(-0.63)	(1.73)	(-0.47)
Tier1$_{t-1}$	0.164	0.463***	0.202	0.683***
	(1.11)	(5.27)	(0.57)	(4.77)
(Dep/TF)$_{t-1}$	-0.0401	-0.0248	-0.0288	-0.0130
	(-0.82)	(-0.74)	(-0.42)	(-0.16)
N	77	57	77	57

t statistics in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

lagged return on assets by 1% is associated with an increase in the probability of receiving positive government support by 41% while an increase in non-performing loans over gross loans by 1% is associated with an increase in the probability of receiving support by 7.4%; all these effects are significant at 1%. The marginal effects are around two to three times smaller for the 2011 case but still significant at 1%. These effects may seem large at first, especially that on profitability. However, it has to be kept in mind that a change in ROAA of 1% is quite big, almost one standard deviation, while a change in non-performing loans over gross loans by 1% is relatively small if compared to its standard deviation which is 4%.

For 2010, conditional on receiving support, a decrease in ROAA by 1% is associated with an increase in liquidity injection over risk weighted assets by 1.5%, around one third of its standard deviation; the effect of an increase of non-performing loans over gross loans by 1% is an increase in liquidity injections over risk weighted assets by about one tenth of its standard deviation.

2.5.3 Cross Validation: Allocation of Credit and Political Influence

Next, we ask whether politically influenced banks tend to facilitate their respective governments in more general terms, not only through purchasing more domestic sovereigns, but also by extending more loans to domestic government institutions. To this purpose we take advantage of the fact that, in 2010 only, the European Banking Authority released data on each bank' s allocation of credit broken down by country of destination and by type of loan; for instance, we know the amount of credit that each bank issued to small and medium enterprises (SME) and to government institutions broken down by the country in which the borrower is located. We then call $DomSME$ the share of domestic SME credit over total SME credit and $DomINST$ the share of loans to public institutions given to the domestic government.

Table 2.7 indeed shows that the effect of political influence of banks' behaviors is not specific to the purchase of domestic sovereign, but it is valid in more general terms: politically controlled banks extend more credit to domestic public entities than other banks do.

Contrary to our expectations, there is not strong evidence that political banks systematically extend more credit to small and medium enterprises; what seems to count to this regard is bank size. This suggests that small banks may proxy for regional banks which tend to lend more locally to small and medium enterprises.

Table 2.7: Allocation of Credit and Political Influence

	(1)	(2)	(3)	(4)
	DomSME_{2010}	DomINST_{2010}	DomSME_{2010}	DomINST_{2010}
$\text{Political}_i/\text{std.dev.}$	4.874^{*}	5.751^{**}		
	(1.87)	(2.06)		
$\mathbb{1}(\text{Political})_i$			8.687	17.17^{**}
			(1.60)	(2.15)
$\mathbb{1}(\text{Cooperative})$	0.483	3.552	0.361	5.232
	(0.04)	(0.37)	(0.03)	(0.54)
$\log(\text{Asset})_{t-1}$	-5.646^{***}	-7.592^{***}	-5.813^{***}	-7.346^{***}
	(-2.90)	(-3.49)	(-2.95)	(-3.46)
$(\text{Dep/TF})_{t-1}$	-0.425^{**}	-0.363^{*}	-0.392^{*}	-0.332^{*}
	(-2.10)	(-1.86)	(-1.86)	(-1.72)
Tier1_{t-1}	0.432	-0.297	0.294	-0.482
	(0.39)	(-0.18)	(0.27)	(-0.29)
N	70	79	70	79

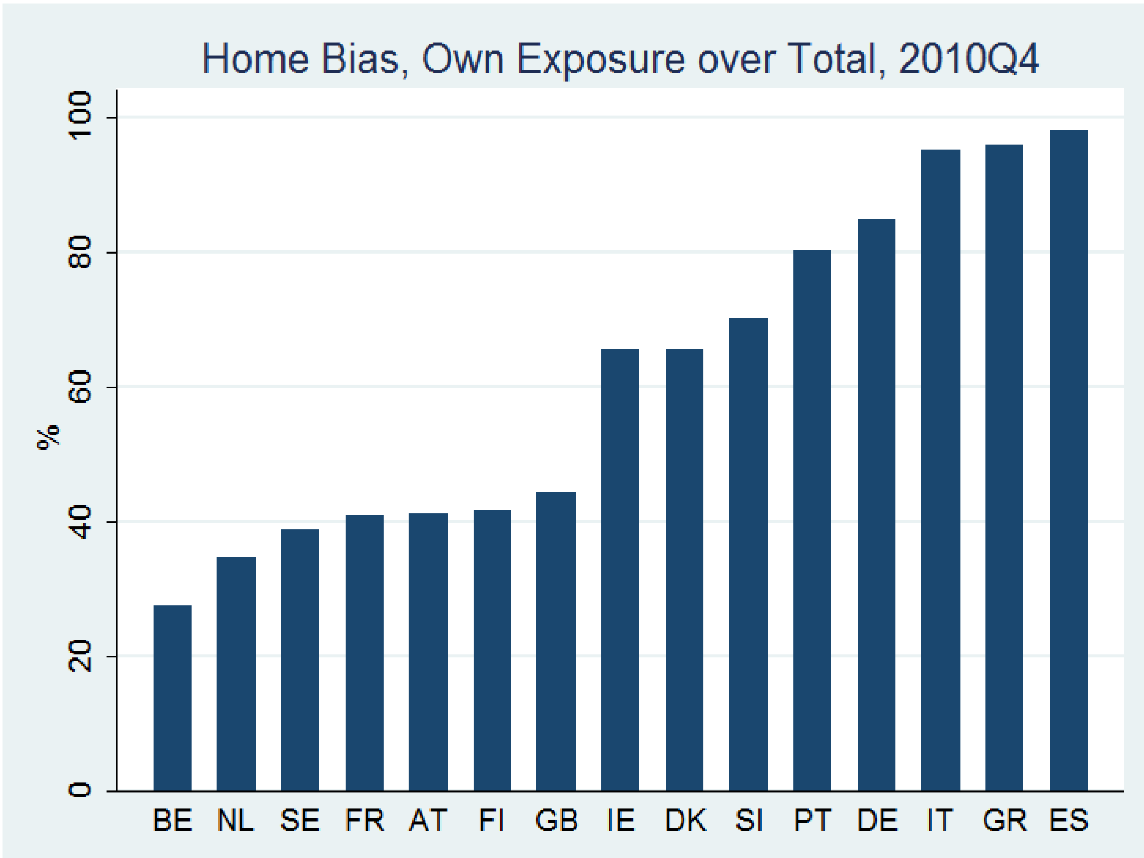
t statistics in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

2.6 Conclusion

We investigate why European banks suffer from a significant home bias in sovereign bond holdings. We believe that certain banks hold a disproportionate amount of their own country's sovereign debt because they are coerced by domestic politicians. In order to test this, we analyze recently collected data from Stress Tests on European banks and we find evidence supportive of this hypothesis: political banks hold more domestic sovereign bonds and they increase their home bias in sovereign holdings conditional on receiving liquidity injections by the respective local governments; this effect is more than twice as big for political banks belonging to the PIIGS than for other European banks. Interestingly, these equity injections seem to be directed towards banks that need it rather than to political banks in particular.

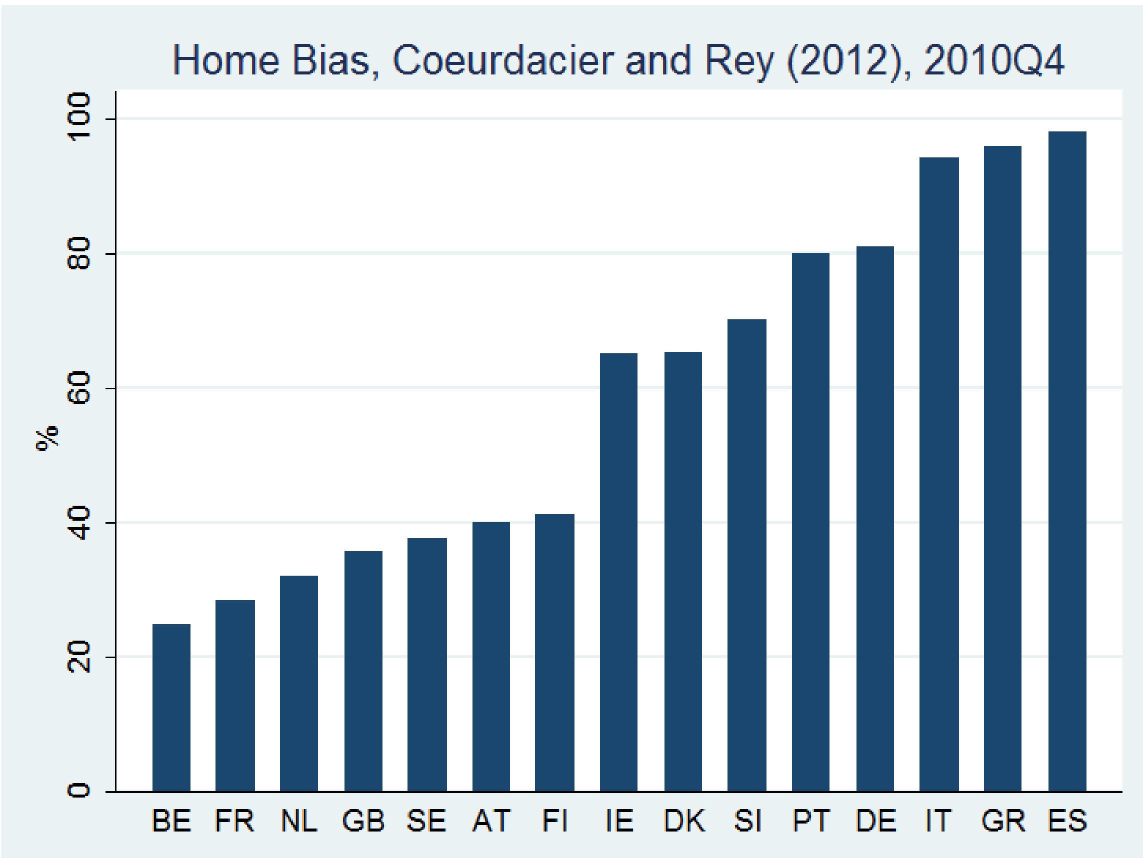
Moreover, we find that politically influenced banks tend to facilitate their respective governments in more general terms, not only through purchasing more domestic sovereigns, but also by extending more loans to domestic government institutions.

Figure 2.1: Median Sovereign Home Bias by country, December 2010



Source: EBA Stress Test 2011. Home Bias defined as the ratio of domestic sovereign by bank b over the total: $HomeBias = Own_b / TotalSovereign_b$. Country codes are the following: Belgium (BE), France (FR), Netherlands (NL), Great Britain (GB), Sweden (SE), Austria (AT), Finland (FI), Ireland (IE), Denmark (DK), Slovenia (SI), Portugal (PT), Germany (DE), Italy (IT), Greece (GR), Spain (ES).

Figure 2.2: Median Sovereign Home Bias by country, December 2010

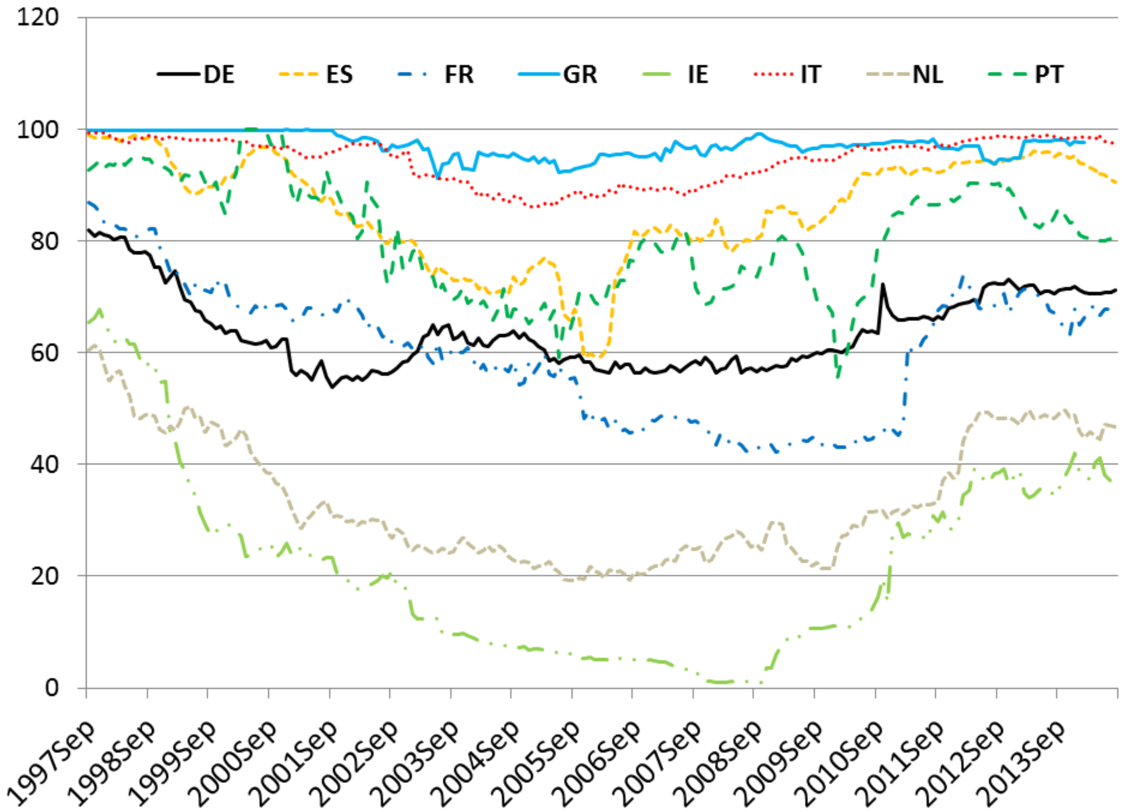


Source: EBA Stress Test 2011. Home bias measure as defined in **?**:

$$HomeBias = 1 - \frac{\text{Share of Foreign Sovereigns in Bank } b \text{ Sovereign Holding}}{\text{Share of Foreign Sovereign Bonds in the Global Portfolio}}$$

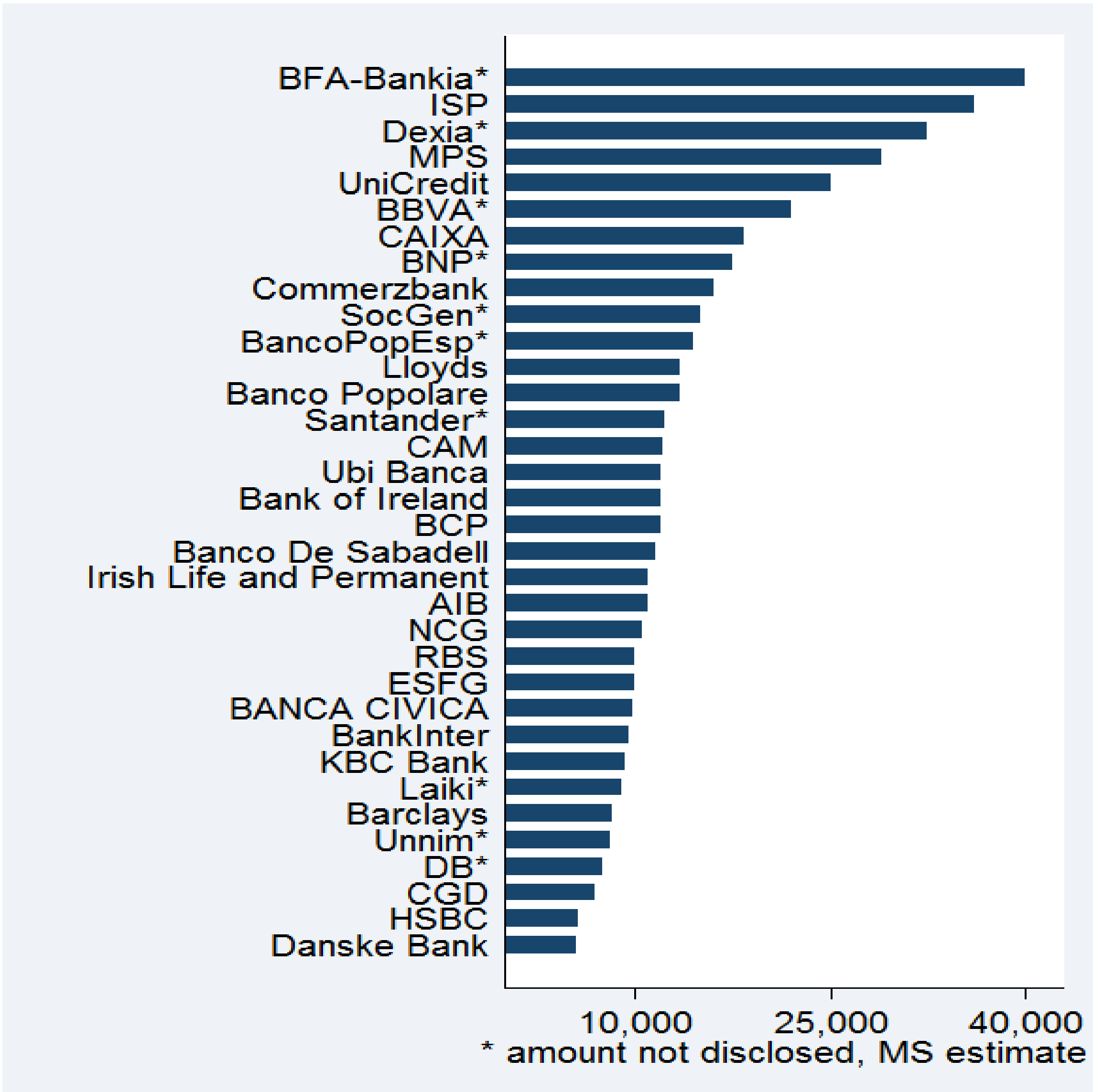
The *Global* portfolio in our case is the EEA30 portfolio, as we have sovereign exposure data for these countries only. When the home bias measure is equal to zero there is perfect diversification; when it is equal to one there is perfect home bias. Anything in excess of zero indicates some level of home bias.

Figure 2.3: Aggregate Sovereign Home Bias by country, Sept 1997 – Sept 2014



Source: ECB Monetary Financial Institutions (MFI) aggregate statistics: ratio between Home and Total of "Securities other than shares" on the Government portfolio (MFI assets). These statistics are given at the country level for all financial institutions (excl. European Central Banks) with a changing composition (*i.e.* this is the raw data that does not take into account mergers&acquisitions and bank failures).

Figure 2.4: LTRO(1+2) 3 year



Chapter 3

The Real Effects of Capital Requirements: Evidence from the UK

3.1 Introduction

Current financial reform proposals focus on the introduction of counter-cyclical capital buffers to maintain financial stability and increase the resilience of financial institutions to adverse shocks. Yet the impact and transmission mechanism of these tools is still not well understood (Galati and Moessner, 2013). This should not be surprising, since Basel III has not been fully implemented yet and in the past, in accordance with Basel I, regulators in most countries imposed a constant capital requirement, by bank and over time. But as extensively documented in Francis and Osborne (2012), Aiyar, Calomiris and Wieladek (2014a,b,c) and Turner (2009), UK regulators adjusted *bank-specific* capital requirements *over time* [1]. In this paper we exploit this unique regulatory regime to test if

[1] In principle, these to address legal, operational and interest rate risks, which were not allowed for in Basel I. In practice, the regulatory decisions on capital requirements for each bank were based on organization structures, IT systems and reporting procedures, rather than financial and balance sheet analysis. Aiyar et al. (2014a,b,c) argue that these institutional characteristics allow to treat changes in capital requirements as exogenous with respect to credit supply.

changes in bank-specific capital requirements affect the asset growth of the SME that borrow from the affected banks. For comparison, we examine the impact of unexpected monetary policy shocks on asset growth and test if these tools reinforce each other. We also study whether certain bank characteristics, such as capital buffers and liquidity, multiple firm-bank relationships or sectoral characteristics on the sources of external finance affect the transmission of this regulatory change.

For capital requirements to be an effective macroprudential policy instrument and affect the aggregate loan supply and have real effects, three conditions need to be satisfied: i) Bank equity needs to be more expensive than bank debt; ii) Capital requirements need to be a binding constraint on a bank's choice of capital and iii) there needs to be only limited substitution from other sources of finance for the borrowing firms. The first condition implies a failure of the Miller-Modigliani (1958) theorem for banks, as otherwise changes in the capital requirement do not need to affect a financial institution's balance sheet. But economic theory provides good reasons for why condition i) should be satisfied, such as asymmetric information (Myers and Majluf, 1984) and the difference in tax treatment between debt and equity [2]. Similarly, empirical work documenting the impact of adverse shocks to capital on loan growth, as in Bernanke (1983) and Peek and Rosengren (1997, 2000) implicitly provides support for this assumption. Regarding condition ii), several empirical studies, namely Ediz et al (1998), Alfon et al (2005), Francis and Osborne (2009), Aiyar, Calomiris and Wieladek (2014a,b,c) and Bridges et al (2014), also demonstrate that capital requirements were a binding constraint on UK banks' capital choices during the 1998-2007 period.Requirements are likely to be especially binding when capital buffers are tight: Banks with small amounts of capital, relative to the requirement, may not be willing to delay the adjustment. This appear to be true in the data: average capital requirements and capital ratios for banks in the bottom quartile of the buffer move almost one-to-one contemporaneously, with a correlation of 0.94, whereas banks in the

[2] See Aiyar, Calomiris and Wieladek (2014a) for an extensive discussion of economic theory and empirical evidence in support of the validity of the first assumption.

top quartile do not appear to be affected by movements in capital requirements (see Figure 1) [3]. This suggests that the second condition is likely to be satisfied as well, especially for banks with low capital buffers.

If conditions i) and ii) hold, the affected bank will have to either raise capital from outside investors, grow capital through retained earnings or cut back on risk-weighted assets, which implies a fall in lending to firms. But the fall in lending need not to have an effect on the borrowing firms if these have alternative sources of funding available (condition iii)). For this reason, in this study we focus mostly on SME firms, which are more likely to be bank dependent. Moreover, we find that our results are stronger for SME with *single-bank* rather than *multiple banking relationships*, as the former cannot easily substitute funding away from the affected bank. Also, we have information on the sources of external finance used by firms at the sector level and hence we can figure out which sectors are most likely to be affected by changes in capital requirements. For example, we know the amount of lending in a sector provided by UK resident banks, which are subject to UK capital regulation, vis–à–vis lending done by branches of foreign banks, which are not regulated by the FSA. Intuitively, the higher the ratio of non-regulated lending to regulated lending in a sector, the smaller the impact of capital requirement changes should be. Our results support this hypothesis.

Several studies use the same underlying FSA capital requirements data to test whether UK bank-specific capital requirements lead to a decline in lending over this period. All of them consider the impact on loan growth to the private non-financial corporate (PNFC henceforth) sector *at the bank level*. Aiyar, Calomiris and Wieladek (2014a), Bridges et al (2014) and Francis and Osborne (2012) find a loan contraction of 5.7%, 5.6%[4] and 5%, following a 1% increase in capital requirements, respectively. There is also related work on *dynamic provisioning*, an alternative macroprudential

[3]This does of course not mean that banks in the top quartile do not adjust at all, but probably with a lag.

[4]To make their results comparable to the other studies, the authors kindly provided us with a figure that refers to the effect over one year and is a weighted average of their results for the commercial real estate and other PNFC lending categories.

instrument, in Spain. In particular, Jimenez et al. (2014a) provide evidence for the impact of dynamic provisioning on lending growth with a detailed loan-level dataset on loans granted and applied for by firms operating in Spain. They find that in good times, 2001Q2-2000Q1, although credit committed declines, there is no change in credit available to firms and hence there is no effect on total assets or employment. They do find some effect on employment growth for the borrowing firms, but only after 2008. Brun et al (2014) use similar credit-registry data from France to examine the impact of the introduction of Basel II in 2007-2008 on French firms and banks. Yet because it coincides with the global financial crisis of 2008, the identification of the real effects of such regulatory change is more difficult.

We examine the effect of changes to bank-specific capital requirements on firm-level asset growth of the borrowers for the period of 1999-2006. In contrast to previous work with this UK dataset, we explore the real economy impact on non-financial firms' balance sheets. As inference in our study is based on many thousands of firm-level observations per financial institution and capital requirement change, econometric bias due to reverse causality is much less likely than in the other studies. More importantly, following the approach in Kwhaja and Mian (2008) and Aiyar, Calomiris, Hooley, Korniyenko and Wieladek (2014), geographical-sector-time dummies allow us to control for loan demand better than in most previous work [5]. Compared to previous work on macroprudential policy tools that use loan-level data, all of our regulatory changes occurred before the global financial crisis. This makes it easier to attribute any effect we find to regulatory changes, as other major bank shocks were absent during our time period. Moreover, we contribute to the discussion on the interaction between monetary policy and prudential regulation. Economic theory provides no clear

[5] The nature of our data does not allow us to use the firm*time fixed-effect pioneered by Kwhaja and Mian (2008), i.e. comparing the lending done to the same firm by differently affected banks (see data section 2.2 for details). However, like in all studies that employ this methodology (Jimenez et al (2012,2014a,2014b), Iyer et al. (2014), Brun et al. (2013) among others), we find that going from an aggregate time effect to a very heterogenous sector-,area-,time effect does not significantly alter any of the baseline results. This suggests that highly disaggregated unobserved heterogeneity is usually not a concern in most partial equilibrium, reduced-form regression settings.

prediction on whether these two policies reinforce or dampen each other (Borio and Zhu (2008), Thakor (1996), Van den Heuvel (2002, 2005)) and empirical evidence on their interaction is scant (Aiyar et al. (2014c)). The UK experience is helpful in this respect, because banking supervision and monetary policy were undertaken by two independent institutions, the FSA and Bank of England respectively. The UK monetary policy did not explicitly take into account capital requirements of individual institutions [6]. It is therefore unlikely that national authorities used capital requirements and monetary policy in conjunction to affect real activity. This means that we can use the narrative measure of UK monetary policy provided in Cloyne and Huertgen (2014) to test if these two policy tools reinforce each other. Finally, we also test if bank characteristics, such as buffers and liquidity, the presence of multiple bank relationships or sectoral characteristics on the dependence on external finance affect the transmission of these regulatory changes.

We find that a rise in an affected banks capital requirement of about 100 basis points, conditional on all observable firm and bank-level characteristics, leads to a decline in asset growth of an SME borrower of about 3.5 to 6.9%, depending on the specification, in the *first year of new bank-firm relationship*. This effect then declines over time [7] . We find a reduction in current liabilities of a similar magnitude, which suggests that the reduction in asset growth is the result of decrease in funding. A 100 basis point unexpected rise in the monetary policy short-term interest rate leads to an asset growth reduction of -4.9%. If both prudential and monetary policy tools are used simultaneously, they amplify each other, but only in the case of a monetary policy expansion. In the case of a tight monetary policy, the two actually attenuate each other. This is consistent with the theory of the *risk-taking channel* of monetary policy (Borio and Zhu (2008)). Moreover, we find that SME borrowing from banks with tight capital buffers are more affected by the regulatory change: this result is very intuitive because banks with tight capital buffers will probably find

[6] The UK also had an inflation-targeting monetary policy regime at that time.

[7] Admittedly, this could also be because the firm has already repaid its debt in the first few years. In fact, our data contain loans made *at origination* and we do not know, in the vast majority of cases, the maturity date.

capital requirements to be binding, whereas banks with larger buffers probably do not. We also find, consistent with evidence from banks' balance sheet in Bridges et al. (2014), that firms in the commercial real estate sector experience the largest reduction. Moreover, SME with multiple (two) bank relationships are not affected by changes in capital requirements, as they can offset the fall in lending from one bank with the other. However, a tightening in monetary policy affects all firms, regardless of the number of their bank relationships. This suggests that while prudential policy tools such as capital requirements affect single-bank firms only, *i.e.* those that cannot easily find alternative sources of external finance, monetary policy shocks can "get in all the cracks". Finally, Aiyar et al. (2014a) find that the presence of competing foreign branches, that are not subject to UK capital regulation, can offset loan contractions associated with greater capital requirements on regulated banks by about 33%. Consistent with their work, we also find that firms in sectors with greater foreign branch presence experience a smaller asset growth reduction of about this magnitude.

Given the lack of previous studies at this highly disaggregated level, it is not straightforward to compare our estimates to those reported in previous work. The closest studies are those by Jimenez et al (2015) and Brun et al (2014). The first study finds *no effect* of capital requirements (dynamic loan provisioning) on firms total assets growth in either good (2001-2000) or bad times (2010-2008) and only *some small* effect on employment growth. Brun et al. (2013) find an elasticity of capital requirements on firms total assets growth ranging from -0.88 to -2.74, but in the crisis period only (2012Q4-2008Q1). But we can compare them to estimates from time-series studies of the aggregate effect of capital requirement shocks on output. Meeks (2014) finds that a 100 basis point rise in capital requirements leads to a peak decline in real GDP of about 1%. How does this compare to our effect? The average fraction of new relationships every year is about 16%, which means a rough translation of the impact reported in this paper is about 7%*0.16=1% of GDP. Similarly for monetary policy, our results are comparable to the impact reported in Cloyne and Hurtgen

(2014). Overall this suggests that our estimates are in line with previous studies. Nevertheless, the effect could still be bigger still, since unlike other micro-level studies, we do not have both loan application and loan outcome data. Quantifying the extent of loan rationing after a rise in capital requirements is therefore not possible. The advantage of our framework is that it is easier to interpret changes in capital requirements, as what they actually are, during a time period without any other major shocks to the banking system. We examine the impact of microprudential capital requirement changes in a partial equilibrium regression framework. The results may of course not be generalizable to macroprudential regulatory changes to capital requirements in a general equilibrium framework. Regardless, in the absence of other evidence, we argue that the estimates provided in this paper could still be informative for policy makers and economic theory.

The remainder of the paper is structured as follows: Section 2 describes the UKs regulatory regime and the data. Section 3 describes the empirical approach for each of the proposed hypotheses, presents the results and examines them for robustness. Section 4 explores the transmission channels in greater detail. Section 5 examines the interaction with monetary policy and Section 6 interactions with other sources of funding. Finally Section 7 concludes.

3.2 UK capital requirement regulation and data

In this section we describe the UK's regulatory regime and the detailed firm-level database that make our investigation possible.

3.2.1 Bank–level Data

Bank-specific capital requirements in most countries were set at a fixed value at or above the minimum of 8 per cent of risk-weighted assets since the introduction of Basel I in 1988. But in the UK, regulators varied bank-specific capital requirements, otherwise known as minimum trigger ratios [8], to address operational, legal or interest rate risks, which were not accounted for in Basel I (Francis and Osborne, 2009). Individual financial institutions were subject to different capital requirements over time and these were subject to review either on an on-going basis or every 18-36 month. This regulatory regime was first implemented by the Bank of England, with the Financial Services Authority (FSA) taking over in 1997. The FSA based regulatory decisions for banks on a system of guidelines called ARROW (Advanced Risk Responsive Operating frameWork), which covers a wide array of criteria related to operational, management, business as well as many other risks. The ARROW approach also encompassed prudential risks, but this was not one of the core supervision areas. Indeed, in his high-level review into UK financial regulation prior to the financial crisis of 2008, Lord Turner, the chief executive of the FSA, concluded that: "Risk Mitigation Programs set out after ARROW reviews therefore tended to focus more on *organisation structures, systems and reporting procedures, than on overall risks in business models*" (Turner, 2009). Similarly, the inquiry into the failure of the British bank Northern Rock concluded that "under ARROW I there was *no requirement on supervisory teams to include any developed financial analysis* in the material provided to ARROW Panels" (FSA, 2008). Hence, one can argue that capital requirement changes are plausibly exogenous with respect to banks' credit conditions [9] Econometric analysis supports the anecdotal evidence. Aiyar, Calomiris and Wieladek (2014a) show that, while bank size and write–offs appear to be important determinants of the *level* of capital requirements in the cross-section – with smaller banks with higher write–offs and lower deposit share having a higher capital requirement – bank balance sheet variables cannot typically predict *quarterly changes* in capital requirements. If

[8] A trigger ratio is the technical term for capital requirement, since regulatory intervention would be triggered if the bank capital to risk-weighted asset ratio fell below this minimum threshold.

[9] Of course, we are not claiming that capital requirement changes are the result of a randomized experiment. we do not need such a strong statement for our analysis to work. All we need is that the regulatory changes are not affected by the behavior of firms borrowing from these banks.

anything, contemporaneous increases in write–off rates predict a *lower* capital requirements [10], but these only explain 1% of the total variation in capital requirements. Moreover, Aiyar, Calomiris and Wieladek (2014b) estimate a panel VAR on PNFC lending and changes in capital requirements at bank balance sheet level. They only find evidence of causality running from changes in capital requirements to PNFC lending growth, but not vice versa. Hence, both anecdotal and econometric evidence therefore suggest that changes in capital requirements were not determined by lending growth and balance sheet characteristics under this regulatory framework. This makes it uniquely suitable to study the impact of changes to capital requirements on the firms that borrow from the affected banks.

The Bank of England has kindly made these regulatory returns data, collected from the BSD3 form, available for our investigation. We collect data on a total of 62 regulated banks' lending to UK PNFCs. Our study covers the time period 1999 to 2006, for two reasons. First, the data after 2006 may have been affected by the start of the UK banking crisis associated with failure of the bank Northern Rock in 2007Q3. Furthermore, prior to 2008Q1, UK regulators relied on the risk-weights associated with Basel I. Unlike the Basel II risk weights, which were adopted after 2008 and can be calculated using banks' Internal Risk Based (IRB) models, these risk weights assigned a weight of 100% to PNFC loans, regardless of individual loan characteristics. This additional regulatory margin would add a further layer of complexity to our analysis. To better isolate how changes in capital requirements affect bank behaviour, we choose the sample period to finish in 2006. Regulated institutions were affected by 100 capital requirement changes during this time, all of which are shown in Figure 2, with summary statistics provided in Table 2. All of the other lender level balance sheet data was provided by the Bank of England's Statistics and Regulatory Data Division [11]. The control

[10] Loan losses (write–offs) are typically recognized on banks' books with some lag, hence the contemporaneous change in write–offs rate correspond to past bad loans. It is then possible that regulators decide to reduce capital requirements at a time when uncertainty about the bank's future losses has been resolved.

[11] All banks operating in the United Kingdom are legally required to provide this information to the Bank of England

variables we derive from these data are described in greater detail in Table 1, with the corresponding summary statistics provided in Table 2.

3.2.2 Firm–level Data

Firm level data come from the Bureau Van Dijk Financial Analysis Made Easy (BvD FAME) database, based on companies' filings with Companies House, the UK's firm registry. The key aspect of this dataset is that it contains the names of the banks (chargeholders) that have secured loans (charges) against each firm. According to Companies House, a charge is defined [12] as the security, such as land, property or financial instruments a company provides as collateral for the loan. While technically the charge is the collateral to the loan, we will use the term as a synonym for the loan itself. The bank-firm relationships in BvD FAME are obtained from the charges registered with Companies House. Charges are legally required to be registered with Companies House twenty-one days after the loan has been created. It is in the interest of the lender to register the charge within the deadline, as otherwise it would not be able to seize the collateral if the company became insolvent. Indeed, the Bank of England surveyed one of the UK's 5 largest lenders in 2013 to examine whether these data actually reflect bank-firm relationships. This was the case for 99.8% of the firm-bank relationships in this dataset, which suggest a high degree of accuracy. Finally, it is important to point out that we only observe the initial charge: Additional funds or re-financing against the same asset later in time does not appear in the dataset. Similarly, we do not observe when firms repay their loans.

There is are some important dimensions in which the data are different from the Credit Registry data used in other work (Jimenez et al. (2012, 2014a, 2014b), Iyer et al. (2014), Brun et al. (2013), Gobbi and Sette (2012)); in particular, they do not contain information on the amount of credit

[12] http://www.companieshouse.gov.uk/about/gbhtml/gp3.shtmlch9

provided by each bank, nor the interest rate charged on each loan. The maturity date of the loan is often missing too. Therefore, we cannot use the firm*time fixed-effects identification pioneered by Khwaja and Mian (2008), as we do not have data on the loan amounts each bank provides to each firm [13]. However, the most important feature of the data is that it allows us to link the banks in our sample to individual firms. Moreover it does so at a specific point of time, which means that we can calculate the length of the bank-firm relationship based on actual transactions rather than on survey data (Petersen and Rajan (1995)). Finally, in orde to examine our results for robustness, the firm-bank panel nature of the data also allows us to include bank-time effects, which fully control for omitted variable bias at the bank, time–varying level

There is a total of 331,000 private, non-financial firms (PNFC) with charges registered in BvD FAME (June 2007 version) and financial data for 1998 to 2006 [14] . We match a total of 252,992 firms to 67 banks with capital requirements data. The matched sample is broadly consistent with the 331k sample of firms with registered charges (see Table 1 in the Appendix). Most of the firms in our sample are small: 92% have assets below £2.8mil, the official threshold for small companies as defined by Companies House [15]. This generates some missing items in the firm-level data, as firms classified as small do not need to report a Profit and Loss (P&L) account and only need to file an abridged version of the balance sheet. Indeed, an examination of the summary statistics in Table 2, reveals that, other than age and the length of the relationship, Total Assets is the most highly populated balance sheet variable, while Turnover and Employees are under-represented. But information on the sector, postcode, date of incorporation and relationship characteristics are

[13] Most of the companies in our sample (88%) have single bank relationship anyways.

[14] Note that a charge may have been registered before 1998 and still show up in BvD FAME. We exclude those charges created before 1990, thus the maximum observed length of a bank-firm relationship in our data is 16 years.

[15] According to Companies House, to be a small company, at least two of the following conditions must be met:

- annual turnover must be £5.6 million or less;
- the balance sheet total must be £2.8 million or less;
- the average number of employees must be 50 or fewer

available for all firms, including small ones. We also have a very good coverage of current liabilities, which include short–term debt, trade credit and taxes, and of credit scores, as these are calculated by an external credit rating agency, CRIF Decisions Ltd, on the basis of both financial and non-financial information (directors' and shareholders' history, County Court Judgements)

Figure 3 shows that a single-bank relationship is dominant, which is not surprising in light of the fact that most of the firms in our sample are small [16] . However, even if only 12% of firms have a banking relationship with multiple institutions, given our large sample size we still have 26,800 firms with two banks and 2,700 firms with three banks. The number of multiple banking firms is therefore sufficiently large to allows us to run separate regressions on these entities. Given their access to multiple banks, we would expect these firms to be less affected by a change in the capital requirement of only one of their relationship banks, relative to single-bank firms, as the other bank can provide a source of credit substitution. The results confirm this expectation. In general, UK firms as whole are less reliant on public debt and equity – and more on bank lending – than the US corporate sector: bank lending represents 65% of total corporate debt in the UK and only 25% in the US (BoE (2011)). This high degree of bank dependence makes the UK's PNFCs relatively more susceptible to bank-level shocks.

Another characteristic of the UK banking system is that it is very concentrated: Figure 4 shows that the top 5 banks provide credit to 91% of firms in our sample. This naturally begs the question of whether we have enough variation in our independent variable of interest, the change in banks' minimum capital requirements, given that most of the sample is dominated by five institutions. One may also wonder if the results are driven by one of these large banks. We show that the results are robust to dropping one of the large banks at a time or all of them at once. Next, Figure 5 shows the

[16] This is different from Braggion and Ongena (2014), who document that UK firms used to have single bank relationship before banking deregulation in 1971 and since then engage in multiple banking. Whereas Braggion and Ongena (2014) focus on listed firms, which tend to be larger and more transparent, we mostly have data on small firms, who are not listed and tend to have single banking relationship.

distribution of firms by sector. The Commercial Real Estate (CRE) sector is the most numerous (37%) followed by Wholesale and Retail Trade (17%), Construction (13%) and Manufacturing (12%). Together these sectors comprise 79% of the companies in our sample. The table below shows the growth rate of firms' total assets broken down by sector. All sectors exhibit positive asset growth on average, consistent with the growth in the UK economy over this period. Among the top five sectors by size, the sectors involved in the UK housing bubble of 2008, the CRE and construction sector, have the higher average growth. However, it is important to notice that even in these sectors, a substantial number of firms (more than 25%) have negative asset growth.

Table 1: Growth Rate of Total Assets by Sector

Sector	Observations	Mean	Std.Dev.	25th	50th	75th
Comm.Real Estate	325,360	0.097	0.461	-0.059	0.028	0.242
Wholesale&Retail	174,746	0.061	0.338	-0.068	0.033	0.171
Manufacturing	141,283	0.05	0.322	-0.086	0.025	0.168
Construction	127,708	0.093	0.437	-0.096	0.065	0.275
Recr.&Pers.Services	72,226	0.059	0.406	-0.088	0.02	0.191
Transportation,Storage,Comm.	38,527	0.076	0.369	-0.079	0.044	0.209
Hotel&Restaurant	34,469	0.064	0.367	-0.055	0.006	0.125
Agric., Hunting, Forestry	17,119	0.033	0.275	-0.051	0.011	0.103
Health and Social Work	13,633	0.128	0.395	-0.033	0.034	0.203
Education	5,843	0.116	0.426	-0.045	0.045	0.225
Mining, Quarrying	1,587	0.114	0.43	-0.057	0.059	0.251
Fishing	1,571	0.034	0.33	-0.074	-0.002	0.112
Public Admin.	567	0.165	0.454	-0.066	0.109	0.347
Utilities	536	0.123	0.453	-0.067	0.061	0.279
Not Available	10,929	-0.058	0.528	-0.131	0	0.042

In conclusion, our dataset contains 252,000 UK PNFC firms, borrowing from 67 different banks with capital requirement changes between 1998-2006. Most firms are small, meaning that they borrow from a single bank, especially one of the big 5, and they are concentrated in the CRE, Wholesale and Retail, Construction and Manufacturing sector. We know both the age of the firm and the length of the relationship with each bank. We have good coverage of firms' credit scores and share of current liabilities over total assets, but not of turnover, profits or number of employees. Finally, the 4digit SIC code and postcode area where these firms operate, allows us to construct sector*area*year fixed effects to control for unobserved heterogeneity among firms.

3.3 Empirical Approach and Results

In this section we describe the empirical framework to test each of the proposed hypotheses and report the results.

3.3.1 Empirical Strategy

In this section we examine whether changes in capital requirements affect the asset growth rate of the borrowing firms with the following regression equation:

$$\Delta \ln Y_{i,t} = \alpha_i + \delta_1 \Delta KR_{j,t} + \delta_2 \Delta KR_{j,t} \frac{1}{1 + L_{i,j,t}} + \phi FC_{i,t} + \gamma BC_{j,t} + T_{h,k,t} + \epsilon_{i,j,t} \tag{3.1}$$

where $\Delta \ln Y_{i,t}$ is the change in the natural logarithm of total assets (or current liabilities) at time t of firm i that took out a loan from bank j. $\Delta KR_{j,t}$ is change in the minimum capital requirement ratio of bank j lending to firm i at time t [17] . $L_{i,j,t}$ is the length of the relationship between firm i and bank j in years [18]. $FC_{i,t}$ is a vector of firm characteristics for firm i at time t. $BC_{j,t}$ is a vector of bank characteristics for bank j at time t, all of which are listed in Table 1. α_i is a lender fixed effect to account for lender unobservable time-invariant characteristics. $T_{h,k,t}$ is a vector of sector-area–year effects to account for unobservable geographical and sector differences, in particular

[17] In the case of multiple bank relationships, this is averaged over the change in capital requirements for all banks.

[18] This variable takes the value of zero in the first year of the relationship, the value of one in the second year of the relationship and so forth.

in loan demand and unobservable firm characteristics, among the approximately 134 areas and 400 sectors that UK firms operate in over time. To avoid perfect multi–collinearity, we drop the first variable from each group of fixed effects. Finally $\epsilon_{i,j,t}$ is assumed to be a normally distributed error term. All standard errors are clustered by firm in each regression specification.

In this empirical model, changes in capital requirements, $\Delta KR_{j,t}$, can affect the growth rate of assets at the firm level, $\Delta \ln Y_{i,t}$, via two different channels: They can affect the dependent variable directly and also through the inverse of the length of the relationship, $1/(1 + L_{i,j,t})$, between firm i and bank j. This specific convex and decreasing functional form for length is chosen so that the effect of capital requirements changes reaches its maximum in the first year of a new bank-firm relationship ($L_{i,j,t} = 0$) and then quickly declines over time, with the curvature allowing for the impact of capital requirement changes to be larger at the beginning of the bank–firm relationship. Economic theory suggests that this could be important for at least two reasons: Gobbi and Sette (2012) and Bolton et al (2014) find that relationship banks attenuate negative shocks to individual firms. This would imply a negative value of δ_2, as the impact of changes to capital requirements, which have a negative elasticity on bank lending, on firms' growth rates would become smaller over time [19]. Finally, there is also a mechanical reason for why it is important to allow a smaller impact over time: firms can repay their loans over time, which we cannot observe, but this would also make the impact on the average asset growth of the borrower smaller [20].

While capital requirements are available at a quarterly frequency, we examine the impact of the change within a year. This is because banks might start to change loan conditions several months before the implementation of regulatory action, due to anticipation, and we cannot observe when

[19] There is also a "dark" side to relationship lending: once the relationship is established, banks extract monopoly rents from their borrowers (Sharpe (1990)). Ioannidou and Ongena (2010), using data on the Bolivian credit registry, find that banks charge a lower interest rate at the beginning of the relatinship and eventually increase it sharply. Our estimate of $\delta_2 < 0$ is consistent with the "positive" view of relationship lending.

[20] The Bank of England surveyed one of the largest five lenders in the UK in 2013 to examine whether their 2012 loan book entries correspond to those in our database. This exercise suggested that about 30 percent of firms paid the debt back within a year. Note that our specification, using a convex and decreasing functional form for length, allow for a much sharper decline after a year, to allow for relationship lending to have an impact as well.

the loan agreement was signed, only when it was recorded. Furthermore, due to adjustment costs, it is quite likely that the reaction to changes in capital requirements occurs with a lag. This is why Aiyar, Calomiris and Wieladek (2014a, 2014b, 2014c) and Aiyar, Calomiris, Hooley, Korniyenko and Wieladek (2014) include the contemporaneous value as well as 3 lags of the capital requirement ratio in their exploration of these data. In other words, even studies that exclusively examine quarterly data allow capital requirements to affect the dependent variable of interest within a year and this is exactly the convention that we follow in this study.

In this study we aim to identify the loan supply effect of changes in capital requirements, and hence it is important to control for loan demand. This is a challenging task in most empirical studies. But the detail of the dataset in this paper allows us to follow the approach presented in Aiyar, Calomiris, Hooley, Korniyenko and Wieladek (2014), who exploit geographical information on the loan destination and use geographical time dummies to control for loan demand. To the extent that loan demand varies over time and across geographical boundaries, this allows them to interpret their estimates as supply effects. Kwhaja and Mian (2008) adopt a very similar approach, except that they observe multiple loans for each firm, unlike in this paper. The information on the sector and location of the firm is very detailed: we know the full 4-digit Standard Industry Classification (SIC) code and the full postcode for all our firms. There is a total of 499 4-digit sectors and 2,757 full postcodes which we group in 122 postcode areas. For example, we can distinguish between farming of cattle in Birmingham and farming of poultry in Coventry. This high level of granularity allows us to construct 4digitSIC*postcodearea*year fixed-effects (150,753 fixed effects) to control as much as possible for sector specific loan demand and compensate for the lack of firm level controls [21]. To the extent that loan demand shocks differ across UK areas and sectors over time, the regression estimates on the change in capital requirements can thus be interpreted as a supply effect.

[21] We will see that the difference between controlling for aggregate, year level fixed effects and the 4digSIC*area*year fixed effect is usually not important in our regressions. This is common with other papers: for example, in Jimenez et al. (2012) the difference between year and firm*year fixed effects is minimal.

Given that equity is expensive and capital requirements are a binding constraint on an individual lender's choice of capital, we expect a decline in the asset growth of a borrower following a rise in capital requirements of the main relationship bank. At firm level, this can happen in two different ways. The affected bank can either reject the application of a borrower that requires a large loan and lend to one with a smaller loan instead or just offer a smaller loan. Unfortunately our dataset does not have information on loan applications, which means that we cannot formally distinguish between these two adjustment channels.

Previous work has examined the impact of changes in capital requirements to PNFC loan supply, but not to the actual asset growth of the firms that borrow from banks affected by capital requirement changes. Francis and Osborne (2012), Aiyar, Calomiris and Wieladek (2014a) and Bridges et al (2014) examine the impact of bank-level changes in capital requirements on lender-level loan growth. An additional advantage of the dataset, compared to previous work, is that it contains many thousand firm-bank observations per capital requirement change. Econometric bias from reserve causality is therefore unlikely. In other words, we argue that equation (1) is substantially better identified than the approaches used in previous work.

3.3.2 Results

Table 4 presents estimates of equation (1) for the growth rate of assets. All specifications include firm fixed effects. Column (1) and (2) only include *year* or *sector* × *area* × *year* fixed-effects, respectively. The coefficient on the interaction term $\Delta KR_{j,t} \times 1/(1 + L_{i,j,t})$ is negative: it implies that the negative effect of a change in capital requirements for the bank(s) j associated to firm i declines as the length of the relationship increases. This is consistent with a "positive" view of relationship lending whereby an affected bank would not cut lending to its long-time customers, but rather offer worse credit conditions to the new customers. Similarly, it is also consistent with the idea that some firms repay their loans within the year, and that hence the average impact of future capital requirement changes will become smaller. Interestingly, coefficients barely change when the aggregate time effect (7 fixed effects) is replaced with a more highly disaggregated $sector \times area \times year$ effect (around 150,000 fixed effects): the stability of the main coefficient of interest is consistent with the findings in other empirical banking papers (Jimenez et al. (2012,2014a,2014b) etc).

One concern with the regressions in columns (1) and (2) is the lack of firm characteristics. In particular, one may worry that our result is not picking up an effect on new borrowers, but rather the fact that banks may decide to cut lending or offer worse credit conditions to young, risky firms. In column (3) and (4) we therefore control for firms credit ratings, the share in current liabilities over total assets (as a proxy for short-term debt exposure) and the age of the firm . These characteristics enter either contemporaneously or lagged. The coefficient on the interaction term is remarkably stable in column (3) compared to columns (1) and (2). Column (4) is our preferred specification in so far as contemporaneous firm characteristics at time t are simultaneously determined with the growth rate of assets at time t and thus more likely to be endogenous than the lagged characteristics. One concern with the regressions in columns (1) and (2) is the lack of firm characteristics. In particular, one may worry that our result is not picking up an effect on new borrowers, but rather the fact that banks may decide to cut lending or offer worse credit conditions to young, risky firms. In column (3) and (4) we therefore control for firms credit ratings, the share in current liabilities over total assets (as a proxy for short-term debt exposure) and the age of the firm . These characteristics enter either contemporaneously or lagged. The coefficient on the interaction term is remarkably stable in column (3) compared to columns (1) and (2). Column (4) is our preferred specification in so far as contemporaneous firm characteristics at time t are simultaneously determined with the growth rate of assets at time t and thus more likely to be endogenous than the lagged characteristics.

Figure 6 plots the partial effect of a change in $\Delta KR_{j,t}$ on the growth rate of total assets

$(\delta_1 + \delta_2/(1 + L_{i,j,t})$) for the estimates in column (4) of Table 4. The negative effect of an increase in capital requirements is greatest during the first year of the impact at -.076 and then dissipates quickly over time, with an effect that is not significantly different from zero by the time 3 years have passed from the beginning of the relationship. As explained previously, this could be either because banks cut back lending by less to firms that they already know, or because firms have already repaid their debt, which could make the average impact much smaller.

Robustness Table 5 explores the robustness of our estimates to both sample selection and modelling choices.

Econometric Assumptions Our presentation of the regression coefficients clearly depends on interpreting $L_{i,j,t}$ as the length of the bank-firm relationship. An alternative interpretation would be that this actually reflects the age of the firm. In that case our regression estimates would imply that older, perhaps safer, firms are less affected by changes to capital requirements. We explore this in the first row of Table 5 by replacing $L_{i,j,t}$ with age in our baseline regression. Now the results are not statistically significant anymore. This supports our interpretation of $L_{i,j,t}$ as the length of the firm-bank relationship. In row (2) we show that our results also seem broadly robust to more aggressive windsorisation, indicating that the results are not driven by outliers.

An important assumption of our regression model is that changes in capital requirements can affect the asset growth of the borrowing firm, even after the relationship has been established ($\Delta KR_{j,t}$ can be different from zero in any year of the bank–firm relationship). Our data does not allow us to observe re-financing and changes in loan terms after the beginning of the relationship, but we wanted to explicitly allow for the possibility that loan terms are adjusted together with changes in capital requirements at any time during the relationship. But this may not be the right assumption to make if loan terms are not frequently re–adjusted. An alternative modeling approach would be to only allow capital requirement changes at the beginning of the relationship ($\Delta\overline{KR}_j$) to affect the asset growth rate of the borrower. In that case, the presence of $\Delta\overline{KR}_j \times 1/(1+L_{i,j,t}$ in the regression means that the initial loan supply impact becomes smaller over time. We re-estimate our baseline model in this way in row (3) of Table 5. The estimates are very similar to the previous results. We also note that this specification is essentially a restrictive version of the baseline model. Because the results do not change, which implies that this restriction does not appear to be binding, we proceed with our baseline model.

Sample Selection: Firms A possible concern related to our results is about the size of the firm in our sample. As it can be clearly seen from the summary statistics in Table 2, our sample is dominated by small firms: the median firm has total assets for £350,000. The risk is that our sample might contain entities which exist purely on paper, but do not actually engage in any economic activity (shell companies). To explore if this is an issue, we exclude the smallest bottom half (£350,000) and three-quarters (£914,000) in rows (4) and (5) of Table 5. In this case the results, if anything, become slightly stronger. To explore whether ownership structure matters, we estimate the baseline regression separately for firms where either more or less than 50% belongs to the same owner in the following two rows ((6) and (7)). This does not seem to make a difference to our results.

Sample Selection: Banks As shown in Figure 4, 91% of the firms in the sample have relationships with 5 large UK banks. The remaining 9% have relationships with 57 different banks. There is clearly a worry that the presence of these large banks may distort our results in one way or another, especially if one believes that large banks, because of their importance, are not subject to frequent changes in capital requirements. To examine if this is the case, we dropped each large bank from the sample in rows (8) through (13). The main coefficient of interest remains very similar throughout and rises only when all of the large banks are excluded at once (row (14)), which is not at all surprising.

Endogeneity: Omitted Variable Bias The main identification assumption in this paper, which we discussed extensively in section 2 is that changes in capital requirements were not determined by the quality of individual banks balance sheets. And although there is anecdotal and econometric evidence to support this assumption, it is plausible that this assumption is not valid. In that case our results could be the result of a third omitted variable. The firm-bank nature of our data allows to us examine if this is an issue. In row (15) of Table 5 we include $bank \times year$ effects in our specification, along with the $sector \times area \times year$ effects as before. The regression estimate on $\Delta KR_{j,t} \times 1/(1+L_{i,j,t})$ is -.051, which is very similar to the baseline regression. Rows (16)–(18) take this even further by allowing the bank–year effects to vary with the sector ($bank \times sector \times year$) or area ($bank \times area \times year$) where the bank is lending or both simultaneously ($bank \times sector \times area \times year$). The results are remarkably stable to the inclusion of all these fixed–effects. Overall this suggests that omitted variable bias at the bank-balance sheet level, and hence endogneity from that source, does not seem to be significant issue in this study, consistent with our main identification assumption.

Overall, Table 5 suggests that the results obtained with the baseline model are robust to sample selection issues at firm and bank level, more aggressive windsorisation, more restrictive econometric modelling choices and omitted variables bias.

Table 6 presents the results of the baseline specification on the growth rate of current liabilities instead of total assets. The results are qualitatively similar and they suggest that the decline in total assets comes from a reduction in short–term liabilities, such as bank debt [22]. Finally, Table 7 includes other bank characteristics such as size (log total assets), liquidity (cash and government bonds over total assets), core funding (retail deposits over total assets), capital buffer and the change in loan write–offs (at time t or $t+1$[23]) both independently (column (1)) and in interaction with

[22]Current liabilities also include trade credit and taxes, so they are a somewhat noisy proxy for short–term bank debt.

[23]The change in loan write-offs are a proxy for the quality of the loan portfolio (UK banks do not report non-

$1/(1+L_{i,j,t}$ (columns (2)–(3)). The coefficient of interest on $\Delta KR_{j,t} \times 1/(1+L_{i,j,t})$ is still significant and if anything gets larger, highlighting the fact that the change in capital requirements we analyze are not driven by specific bank balance sheet characteristics.

3.4 Transmission Channels

Capital requirements will only affect firms through the loan supply channel only if they are binding for. If banks were subject to a higher capital requirement, but had large capital buffers to start with, they may always just choose to hold a smaller buffer. Banks with very tight capital buffers, on the other hand, do not have this luxury: they have to bear the brunt of the adjustment through a reduction in risk-weighted assets. This is confirmed visually by inspecting Figure 1: capital requirements and capital ratios for banks in the bottom quartile move almost one–to–one and have a coefficient of correlation of 0.94, suggesting that banks with low buffers probably find capital requirements changes to be binding. However, banks in the top quartile do not seem to be affected by changes in capital requirements. Similarly, banks with a lot of liquid assets, such as cash and government bonds, should be less affected by changes in capital requirements. In fact, although these assets carry a risk-weight of zero and thus selling them will not directly affect risk-weighted assets, if they are sold at profits they can increase retained earnings and therefore allow to rebuild capital organically.

We explore these hypotheses, by interacting the main coefficients of interest with a dummy variable that takes the value of one if a bank's capital buffer or liquidity ratio is in the top or bottom quartile of the distribution, and zero otherwise, in Table 8. The bottom panel shows the partial derivative of total asset growth with respect to changes in capital requirements for different values of these dummy

performing loans). Since loan losses at time t usually refer to loans from year t-1, we introduce the lead change in write-offs as an alternative control.

variables. Column (1) suggests that banks with high capital buffers are not affected by changes to capital requirements. On the other hand, banks in the middle (lower) quartiles of the distribution are more (most) affected, consistent with the economic theory. A similar pattern applies to the corresponding dummy variables for liquid assets: banks with high (medium) [low] liquidity ratios are not (more) [most] affected. These effects are robust when interaction terms for both capital and liquidity, are included jointly in column (3).

Tables 9 and 10 explore whether our baseline effects vary by either firm characteristic or sector in an economically meaningful way. Table 9 shows estimates of the baseline model for different economic sectors separately. This exercise suggests that the effects are strongest in the commercial real estate sector. This is consistent with the evidence in Bridges et al. (2014) that use bank balance sheet data and find the largest impact on the loan supply in the commercial real estate sector. In Table 10, we present estimates for a regression model where the main coefficients of interest have been interacted with dummy variables in a similar fashion to Table 8, but using firm rather than bank characteristics. In particular, these dummies that take the value of 1 if firm age (column (1), the credit rating (column (2)) or the current liability to total asset ratio (column (3)) or all together (column (4)) are in either the top or bottom quartiles of their respective distributions. These do not suggest a significant degree of variation by individual firm characteristics, as opposed to the strong results by bank characteristics.

3.5 Do Monetary and capital requirement policy reinforce each other?

A recent issue of interest in central banking is whether monetary and capital requirement policy should be co-ordinated. Clearly if each instrument has one target, and the effects of both instruments are completely orthogonal, no co-ordination is necessary. In that world monetary policy would only focus on price stability, while capital requirement policy would only address financial stability. However, if one instrument affects the transmission of the other, then this will need to be taken into account by the corresponding policy committee to avoid under or overshooting of the target, which can have socially undesirable consequences. Whether co-ordination is desirable or not therefore depends on whether these instruments reinforce the effects of each other. Economic theory suggests that these two instruments should affect the transmission of each other, but the direction of the effect is unclear. The theory has explored the extent to which these two instruments affect the transmission of each other through three different channels: the *bank capital channel*, the *risk-taking channel* and the *term-structure channel* of monetary policy.

The *bank capital channel* of monetary policy predicts that monetary policy and changes in capital requirements reinforce the effects of each other. Van den Heuvel (2002) shows that an unexpected monetary policy contraction can lead to smaller capital buffer, as a result of realised interest rate risk. This happens because of the maturity mismatch on banksbalance sheets between assets (*long duration*) and liabilities (*short duration*), so that an increase in the interest rate causes profits and hence capital to decline. This means that a coincident rise in capital requirements will have a larger impact on the loan supply, as it is likely to be more binding.

The *risk-taking channel* of monetary policy suggests instead that the sign of the interaction may be asymmetric, depending on the sign of monetary policy. In an environment where banks target a fixed nominal return, a monetary policy expansion and the associated reduction in interest rates may lead to a search for yield and a rise in bank leverage (Borio and Zhu (2008), Adrian and Shin (2011)). Empirical evidence from Spain (Jimenez et al. (2014b)) and Bolivia (Ioannidou et al. (2014)) points out that a lower overnight rate induces lowly capitalized banks to take on more risk than highly capitalized banks, where the risk is measured with the presence of a bad credit history with nonperforming loans and a higher subsequent probability of default. This may further lead banks to reduce capital buffers. Given that banks with tight capital buffers are more likely to cut back risk-weighted assets, a change in capital requirements will have a greater impact on loan supply in this situation. On the other hand, during periods of monetary policy tightening, it is not clear how capital buffers would respond. In that situation a rise in capital requirement may be less binding on the actual capital ratio and therefore have a smaller impact on the loan supply. In other words, it is plausible that the sign of the interaction between these two instruments depends on the sign of the monetary policy action.

Finally, Thakor (1996) argues that the sign of the interaction between these two instruments will depend on the reaction of the *term structure* of interest rates. If, following a monetary expansion, long rates fall by more (less) than short rates, implying a decrease (increase) in the interest rate term premium, long–term government securities will become less (more) profitable, compared to lending. Since in the presence of risk–based capital requirements government securities have a zero–risk weight, the incentive to shift to lending (bonds) increases. In that case, a coincident decline in capital requirements will lead to a smaller (greater) impact on the loan supply. Whether or not monetary policy reinforces or counteracts changes in capital requirements therefore depends on the yield curve reaction.

Despite this rich body of economic theory, empirical work that attempts to test these different transmission mechanisms is still scarce. Aiyar, Calomiris and Wieladek (2014c) are one of the first studies to undertake this task with UK bank-level balance sheet data. Across a large variety of many different specifications, they do not find any statistical evidence that these two tools reinforce each other. In this paper, we repeat this exercise, but with two important differences. We have more granular data and a more exogenous measure of UK monetary policy. Specifically, we use the measure proposed in Cloyne and Huertgen (2014), which is the equivalent of the Romer and Romer

(2004) series of exogenous monetary policy shocks for the UK. From an econometric modelling perspective, we assume that monetary policy affects asset growth of the borrowing firms through exactly the same channels as changes in capital requirements. That is we again assume that firms will be less affected by monetary policy changes as the bank-firm relationship becomes longer over time.

The results from this specification are shown in Table 11. Column (1) shows that once monetary policy is included, a 100 basis points rise in monetary policy leads to an asset growth contraction of about 4.9%, slightly smaller than the 6.9% contraction following a 100 basis points rise in the capital requirement ratio. Column (2) adds the interactions between the monetary policy and capital requirement terms. Neither $\Delta KR_{j,t} \times MPshock_t$ nor $\Delta KR_{j,t} \times MPshock_t \times 1/(1 + L_{i,j,t})$ have a significant effect on the growth rate of assets. This suggests that, at first sight, there is no interaction as in Aiyar, Calomiris and Wieladek (2014c) and contrary to the bank capital channel in Van den Heuvel (2002). Column (3) introduces a triple (and quadruple) interaction term of capital requirements and monetary policy shocks with annual changes in the term premium (and $1/(1 + L_{i,j,t})$) to examine to which extent the term premium channel (Thakor (1996)) operates. All the interaction terms between the monetary policy shock and capital requirements changes are not significant, suggesting that the term structure does not matter in this case. Column (4) allows the interactions to vary with the sign of the monetary policy surprise, by interacting the corresponding coefficients with a dummy variable ($TIGHT_t$) that takes the value of one during a monetary policy tightening and zero otherwise. First of all, notice that the monetary policy shock has an independent and significant negative impact on the growth rate of assets only for a monetary policy tightening (-12.8%) and not for a loosening. Further, all the interaction terms (triple or quadruple) between capital requirements and the monetary policy shock are now significant. In particular, they indicate that there is an additional negative effect of monetary policy and capital requirements in the case of a monetary policy expansion (+0.0364-0.157=-12%), so that the two policies reinforce each other. During a monetary policy contraction, on the other hand, the sign of the monetary policy capital requirement interaction switches (+0.0364-0.157-0.077+0.348=+15%), so that the two policies attenuate each other if used simultaneously. Overall, these results are consistent with the risk-taking channel of monetary policy, that suggests that the two policies reinforce each other only in the case of a monetary policy expansion.

3.6 Credit Substitution: Multiple Relationship and Other Sources of Funding

So far we have examined to which extent changes in capital requirements on individual banks affect the asset growth rate of the borrowing firms within a partial equilibrium framework. But in general equilibrium, alternative sources of credit, such as other banks, capital markets or trade credit, which are not affected by capital requirement changes, may offset these effects. As a result, it is not clear if micro-prudential policy actually affects the loan supply. Previous work has used bank-level data to test for the presence of such credit substitution for PNFC firms operating in the UK from 1999 -2006. Aiyar, Calomiris and Wieladek (2014a) and Aiyar, Calomiris and Wieladek (2014b) find evidence for credit substitution from foreign branches, but less from capital markets.

In Table 14, we interact the fraction of either foreign branch lending ($BLnonreg$), that are not suject to UK capital requirements, or capital markets finance (equity + bond issuance, $Nonbank$) over the lending done by UK residents bank ($BLreg$), that are subject to capital requirements, at the sector level with our main coefficients of interests. The bottom panel of Table 14 shows the partial derivative of total asset growth with respect to changes in capital requirements for a zero or one standard deviation increase in either $BLnonreg/BLreg$ or $Nonbank/BLreg$. We can see that for firms with a one standard deviation increase in lending by foreign branches over regulated lending the effects of capital requirements is attenuated by 1.5% without including lagged firm characteristics (column (1)) and by 3% including these (column (3)). Increasing capital markets finance over regulated lending (columns (2) and (4)) has no significant effect instead. These regression results overall confirm the previous findings in Aiyar, Calomiris and Wieladek (2014b) that credit substitution from foreign branches, in these data, is substantially stronger than that from capital markets.

However, given that we analyse bank-specific capital requirement changes and that these are mostly uncorrelated in the cross-section, our data allow for a much more powerful test of the credit substitution hypothesis. If financial entities, which were not affected by changes in capital requirements, truly are a source of credit substitution, then firms with multiple bank relationships should not be affected by capital requirement changes to only one of their relationship banks, as these firms can substitute funding away from the affected bank. In regression model (1), $\Delta KR_{j,t}$ is the change in the capital requirement of bank j at time t for single bank-firm relationships. For multiple bank-firm relationships, $\Delta KR_{j,t}$ is the average of the capital requirement changes of all banks related to one particular firm. In this section we allow the capital requirement of each bank to affect the asset growth of the related firm individually, rather than as an average. We can only undertake this exercise for those firms which have exactly two relationship banks, as there are too few banks with three relationships banks or more. The results from this exercise are presented in Table 12, either without firm controls (column (1)), with contemporaneous controls (column (2)) or lagged (column (3)). That clearly suggests that for a firm with two relationship banks, changes in capital requirements do not affect asset growth, as they do instead for firms with single banking relationship (column (4)).

Table 13 repeats this exercise for monetary policy. Estimated on the same sample, monetary policy affects all of the firms through the length of the relationship, regardless whether the firm is a single or multiple relationship bank. This is consistent with the idea that since monetary policy is an aggregate policy instrument, it affects all firms and banks regardless of whether they have access to alternative sources of credit or not. This is a novel result in the empirical literature and it carries important policy implications for the use of monetary versus prudential policy.

3.7 Conclusion

Countries around the world have introduced macroprudential regulation to increase the resilience of the financial system to socially costly financial crises. One proposed instrument, which is also embedded in Basel III, is a time-varying capital requirement. But, to date, there is only little understanding of how this instrument will affect the real economy. The UK's unique regulatory regime, where banks and building societies were subject to time-varying capital requirements, together with a new firm-bank level database, covering all reporting real economy firms in the UK between 1998 to 2006, allows us to provide a first empirical examination of this important question. The purpose of this paper is to examine the impact of changes to capital requirements on the asset growth rate of the borrowing firms. UK Banks were subject to time-varying capital requirements, which varied by institution and over time. Economic theory suggests that if an increase in capital requirements is binding and the Miller-Modgliani (1958) theorem fails, then the affected institution will either need to raise capital or reduce risk-weighted assets to satisfy the new requirement. Previous work, such as Aiyar, Calomiris and Wieladek (2014a), Francis and Osborne (2009) or Bridges et al (2014) tests this last implication on balance sheet level PNFC lending. In this paper we examine whether this loan supply effect carries through the to the asset growth rate of individual firms. We also compare our effect to monetary policy and examine to which extent credit substitution can offset some of these effects. Our results suggest that that a rise in a banks capital requirement of about

100 basis points leads to a decline in the asset growth rate of borrowing firm of about 3.9% to 6.9%. A 100 basis point unexpected rise in the monetary policy short-term interest rate leads to an asset growth reduction of -4.9%. If both tools are used simultaneously, they amplify each other, but only in the case of a monetary policy expansion. SME borrowing from banks with tight (loose) capital buffers are more (less) affected by the regulatory change. Consistent with evidence from banks' balance sheet in Bridges et al. (2014), firms in the commercial real estate sector experience the largest reduction. SME with multiple bank relationships are not affected by changes in capital requirements to only one of their relationship banks. But a tightening in monetary policy affects all firms, regardless of the number of their bank relationships. This suggests that while prudential policy tools such as capital requirements affect single-bank firms only, i.e. those that cannot easily find alternative sources of external finance, monetary policy shocks "get in all the cracks". Finally, Aiyar, Calomiris and Wieladek (2014a) find that the presence of competing foreign branches, that are not subject to UK capital regulation, can offset loan contractions associated with greater capital requirements on regulated banks by about 33%. Consistent with their work, we find that firms in sectors with greater foreign branch presence experience a smaller asset growth reduction of about this magnitude.

APPENDIX – Data

BvD FAME data come from the June 2007 CD-ROM. The search strategy filters firms according to the following criteria:

- Exclude Financial and Insurance companies (2003 SIC codes from 6500 to 7000)
- At least one year of Total Assets
- At least one registered charge

This yields a total of 331,996 companies operating in the UK. Note that the search strategy deliberately does not consider filters on turnover, employees or other variables, as this would create a reporting bias in favour of medium and large companies.
We match a total of around 252k firms to 67 banks for which we have capital requirements data. Notice that we do not match all of the 331k firms because either the *chargeholder* is a non-bank (private citizens, finance companies or other funds) or because the c*hargeholder* is a branch of foreign bank, not subject to UK capital regulation.
In a few cases the name of a bank is listed under *chargeholder* for a charge with a firm although the bank is only acting as an agent for another lender. Large banks often times act as agents for another lender to monitor and screen the borrower in exchange for a fee. Therefore, it may not be correct to match a firm to these agent-acting banks, as capital requirements changes at the agents' level should not affect credit conditions for the borrower (although sometimes the bank is listed as "Agent acting for itself"). Other such non clear roles are "security trustee", "agent trustee". There are only 2,581 firm-year observations with such unclear roles: we can safely exclude them from the regressions and the results are not affected.

Table 1 - Appendix: Comparison between matched and non-matched sample

Variable	Observations	Mean	10^{th}	50^{th}	90^{th}
331k non matched sample					
Total Assets	1,921,170	3.56	.02	.26	1.9
Turnover	471,825	8.7	.02	.37	8.3
Employees	185,720	183.6	3	37	234
252k matched sample					
Total Assets	1,1485,854	2.47	.053	.35	2.37
Turnover	286,789	7.69	0.036	0.52	10.8
Employees	126,575	156	3	45	243

Table 1A shows some summary statistics for the two samples, the 331k non matched sample and the 252k matched sample. Rather than looking at the means, which are not robust to outliers, we can look at the other percentiles of the distribution to see that the two samples are broadly consistent across the two

FIGURES

Figure 1: Capital Requirements and Capital Ratios by Bank Buffer

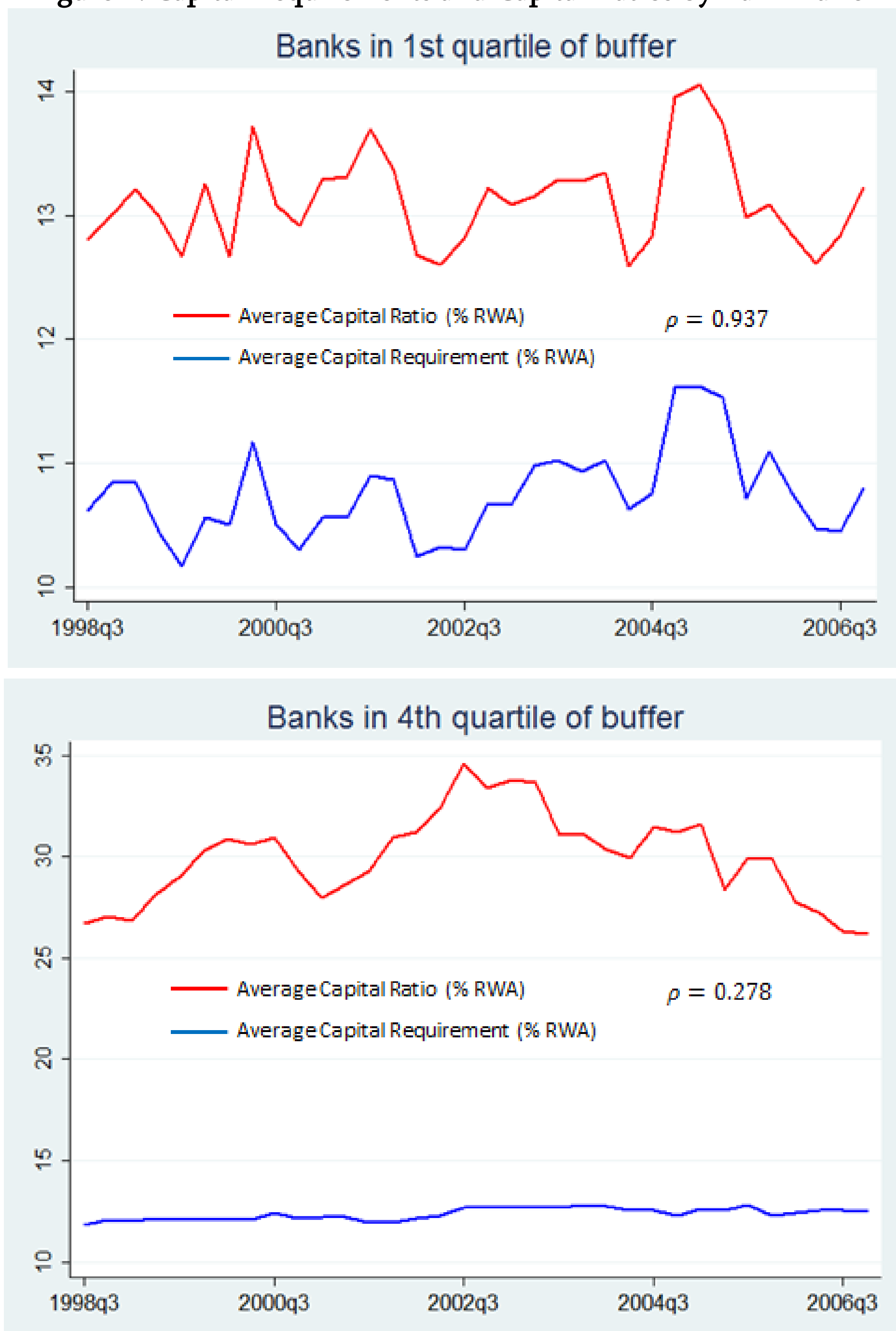


Figure 2: Number of Capital Requirement Changes

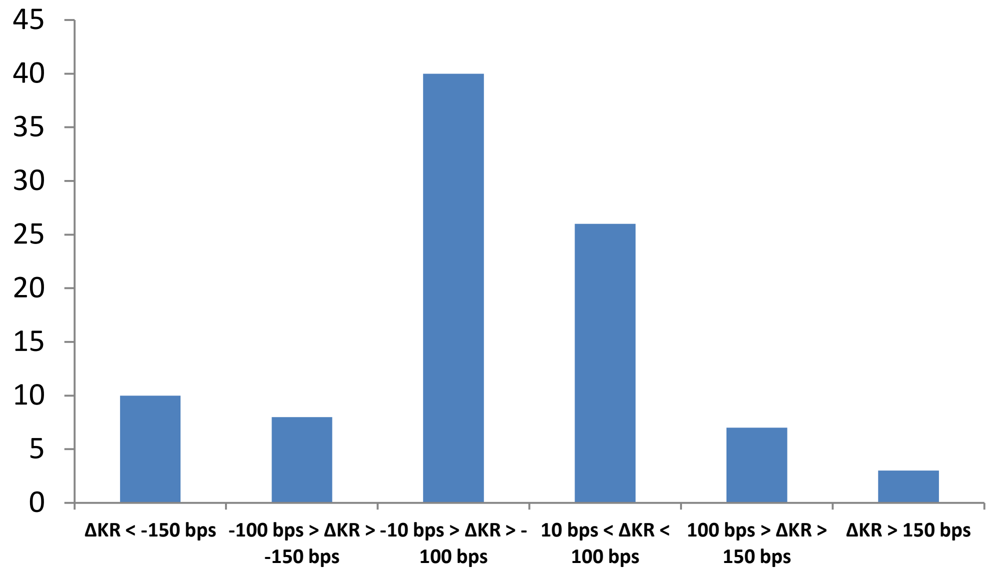


Figure 3: Number of Banks by Firm

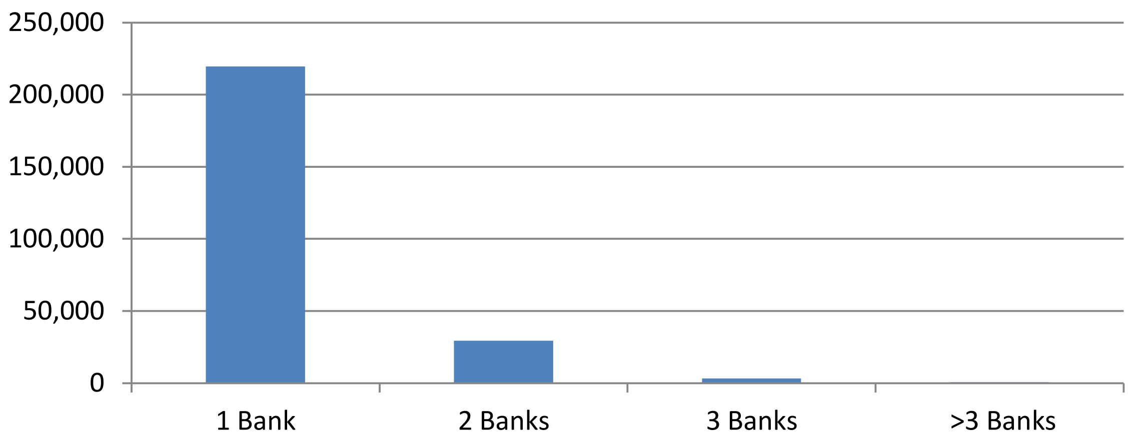


Figure 4: Market Shares by Bank

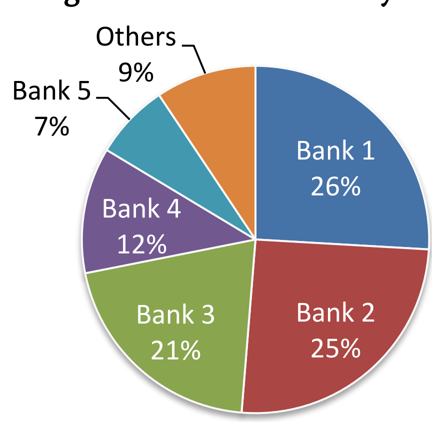


Figure 5: Number and Percentage of Firms by Sector

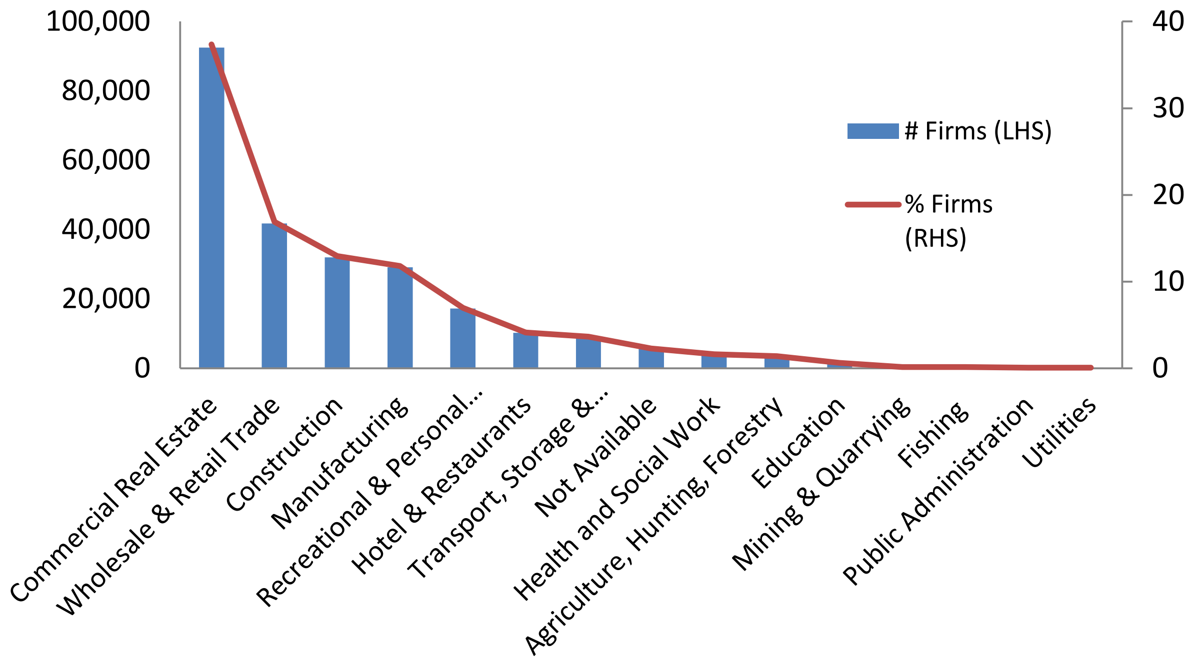
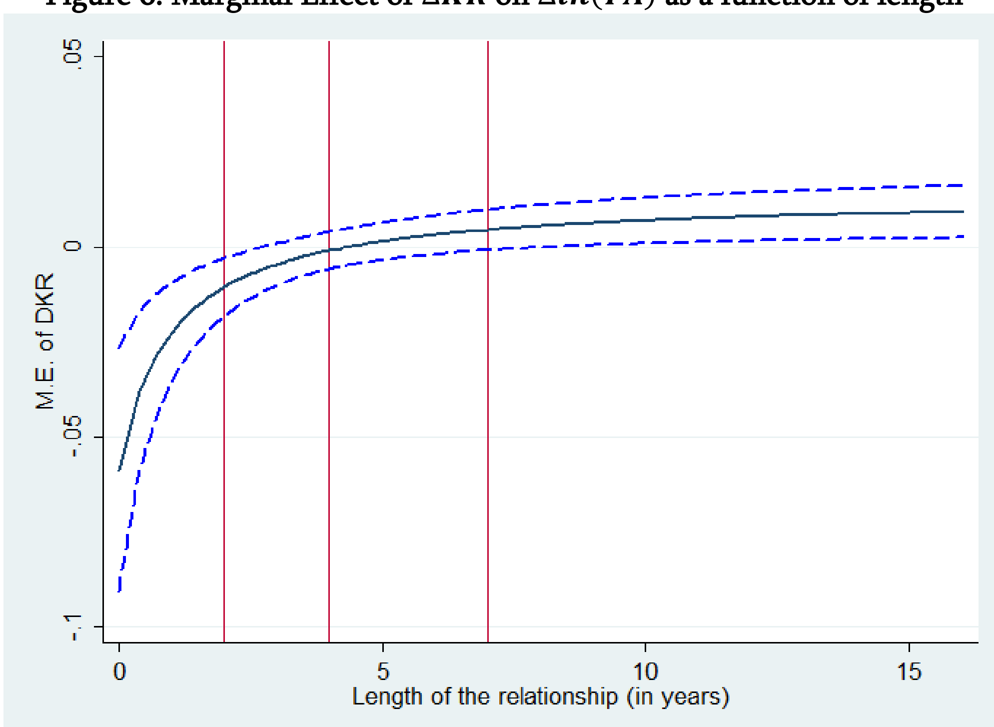


Figure 6: Marginal Effect of $\Delta \boldsymbol{KR}$ on $\Delta \boldsymbol{ln}(\boldsymbol{TA})$ as a function of length



This figure contains the plot of the marginal effect of a change in capital requirement on the growth rate of assets for the baseline specification including lagged firm controls. The partial derivative is a function of length and the vertical red lines represent the 25th (2years), median (4) and 75th (7) percentiles of the distribution of the firm-bank relationship length in years. Dashed lines show the 95% confidence interval around the point estimate.

TABLES

Table 1: Variable Description		
Variable	**Definition**	**Notes**
Firm and bank-firm relationship data (BvD FAME)		
Δln(TA)	Growth rate of Total Assets	
CredScore	Credit Score (QuiScore)	From 0 (worst credit score) to 100 (best). Calculated by CRIF Decision Solutions Ltd., see Table 1a for details.
Currliab/TA	Current Liabilities over Total Assets ratio	Current liabilities include: Short Term Loans & Overdraft, Trade Creditors and Taxes & Dividends
Age	Years since date of incorporation	
Length	Years since creation of the loan	
BvD Independence Indicator	Description of ownership structure	From A (no shareholder with more than 25% ownership) to D (one shareholder with >=50%)
Turnover	Turnover (Sales) in £ mil.	
Employees	Number of Employees	
Bank level data (BSD3 and QFS forms)		
KR_t	Minimum ratio for capital-to-risk weighted assets (RWA) for the banking book.	The BSD3 form provides this information for Banks. QFS provides it for Building societies.
ΔKR_t	Yearly change in KR	
$\Delta Woff_t$	Yearly change in writeoffs rate: nbpa550t/nhd510	nbpa550t – Loan writeoffs nhd510 – Total Risk Weighted Assets
Bank Size	Natural log of (BT40)	BT40 – Total Assets
Bank Liquidity	Liquid to Total Assets ratio: (BT21+ BT32D)/BT40	BT21 - Cash ; BT32D – Holdings of Government Stock
Core Funding	Deposits to Total Assets ratio: (BT2H +BT3H)/BT40	BT2H – Retail Sight Deposits, BT3H – Retail Time Deposits
Buffer	Actual Capital Ratio (nhd40/nhd510) - KR	nhd40 – Total Eligible Capital, nhd510 - RWA
Sector level data on external finance		
BLreg	Regulated Lending	Lending by UK resident banks and UK subsidiaries of foreign banks

BLnonreg	Non-regulated Lending	Lending by UK branches of foreign banks
Nonbank	Non bank external finance	Sum of equity and corporate bonds outstanding

Table 2: Summary Statistics

Variable	Observations	Mean	Std.	25th	50th	75th
Firm and bank-firm relationship data (firm-year panel BvD FAME)						
Total Assets(£mil.)	1,146,711	2.12	53.75	0.137	0.351	0.914
Δln(TA)	967,551	0.076	0.406	-0.072	0.029	0.204
Turnover (£mil.)	285,434	4.54	13.42	0.143	0.522	2.431
Employees (#heads)	125,453	146.5	1049.6	11	44	103
CredScore	1,107,154	50.2	22.60	35	48	64
Currliab/TA	1,139,230	0.63	0.553	0.307	0.547	0.809
Age	1,413,935	12.57	13.68	4	8	16
Length	1413944	4.82	3.922	2	4	7
Bank level data (bank-year panel)						
KR_t (%)	520	11.55	3.07	9	10	13
ΔKR_t (bps.)	520	-4.1	57.2	0	0	0
$\Delta Woff_t$ (bps.)	516	1.655	97.5	-3.972	0	6.365
Size (in £mil.)	520	32,590	89,367	710	4,104	15,100
Liquidity (bps)	520	161	271	1	43	220
Core Funding (%)	520	51.04	30.79	21.92	53.52	79.57
Buffer (%)	517	16.58	33.32	3.1	6.8	14.4
External Finance Dependence data (sector-year panel)						
BLnonreg/BLreg	126	0.52	0.65	0.04	0.26	1.18
Nonbank/BLreg	126	1.29	2.7	0.006	0.2	4.4

Table 4 – Capital Requirements and Firms' Growth Rate of Assets

	(1)	(2)	(3)	(4)
$\Delta KR_{i,t}$	0.008***	0.007**	0.009**	0.013***
	(0.00266)	(0.00368)	(0.00367)	(0.00437)
$1/(1+L_{i,t})$	0.425***	0.425***	0.412***	0.459***
	(0.00411)	(0.00571)	(0.00590)	(0.0113)
$\Delta KR_{i,t} \times 1/(1+L_{i,t})$	-0.0340***	-0.0352***	-0.0369***	-0.0721***
	(0.00946)	(0.0130)	(0.0131)	(0.0197)
$CredScore_{i,t}$			0.001***	
			(5e-5)	
$Currliab/TA_{i,t}$			-0.154***	
			(0.004)	
$CredScore_{i,t-1}$				0.001***
				(4.6e-5)
$Currliab/TA_{i,t-1}$				0.130***
				(0.005)
Young			0.03***	0.018***
			(0.003)	(0.002)
Old			0.007**	0.003
			(0.003)	(0.003)
Firm Fixed Effects	X	X	X	X
Year Fixed Effects	X			
SIC4dig×Postcode ×Year Fixed Effects		X	X	X
Observations	969,052	968,012	927,310	871,423
N of firms	212,894	212,894	208,316	196,980
R squared	0.338	0.444	0.462	0.445

This table presents the results for the baseline regression. The dependent variable is the log difference of total assets by firm i between time t and t-1. All specifications include firm fixed-effects and either year fixed effects (column (1)) or SIC4dig*postcodearea*year fixed-effects (columns (2)-(5)). $\Delta KR_{i,t}$ is the change in capital requirements between year t and year t-1 for the bank lending to firm i (it is averaged over all banks lending to firm i at time t in case of multiple banks). $L_{i,t}$ is the length of the relationship between firm i and its banks, measured in years since the creation of the loan. $CredScore_{i,t}$ is the Credit Score (QuiScore) of the firm, measured on a scale of 0 (worst risk) to 100 (no risk). $Currliab/TA_{i,t}$ is the ratio of Current Liabilities (short term debt, trade credit, taxes and dividends) over total assets. $Young, Old$ are dummies for the age of the firm, at the 25th and 75th percentiles (4 and 16 years old respectively). Finally, $\Delta\ln(TA)_{i,t-1}$ is the lagged growth rate of total assets (lagged dependent variable). For statistical significance, we use the following convention throughout: *** p<0.01, ** p<0.05, * p<0.1. Firm-clustered standard errors are reported in parenthesis.

Table 5: Robustness of Baseline Results

Robustness Exercise	$\Delta KR_{i,t}$	$1/(1+L_{i,t})$	$\Delta KR_{i,t} \times 1/(1+L_{i,t})$
Econometric Assumptions			
(1) Replace $L_{i,t}$ by firm age	0.00415	0.739***	-0.0430
(2) Dependent variable winsorised at 5%&95%	0.007**	0.321***	-0.038***
(3) Time-invariant capital requirement change		0.478***	-0.0715***
Sample Selection – Firms			
(4) Exclude small firms (size in bottom 50%)	0.015***	0.502***	-0.093***
(5) Exclude small firms (size in bottom 75%)	0.019**	0.504***	-0.116***
(6) Exclude firms with < 50% single ownership	0.0146**	0.436***	-0.0654**
(7) Exclude firms with > 50% single ownership	0.0151**	0.500***	-0.0835**
Sample Selection – Banks			
(8) Excluding Barclays	0.013***	0.469***	-0.07***
(9) Excluding HSBC	0.014***	0.461***	-0.071***
(10) Excluding NatWest	0.017***	0.474***	-0.098***
(11) Excluding Lloyds	0.015***	0.468***	-0.086***
(12) Excluding RBS	0.013***	0.457***	-0.077***
(13) Excluding HBOS	0.012***	0.445***	-0.072***
(14) Excluding all Big banks together	0.025***	0.514***	-0.123***
Endogeneity - Omitted Variable Bias			
(15) Include Bank-year effects	-0.0185*	0.471***	-0.0508***
(16) Include Bank-Sector-year effects	-0.0191*	0.465***	-0.0586***
(17) Include Bank-Area-year effects	-0.0144	0.467***	-0.0438***
(18) Include Bank-Sector-Area-year effects	-0.0320**	0.493***	-0.0590**

The dependent variable is the log difference of total assets by firm i between time t and t-1. All specifications include firm fixed-effects and SIC4dig*postcodearea*year fixed-effects, unless otherwise stated. Lagged firm variables, not shown are: $CredScore_{i,t-1}$, $Currliab/TA_{i,t-1}$ and $young, old$ Standard errors are clustered by firm.

Table 6: Impact of Capital Requirements on Current Liabilities

	(1)	(2)	(3)	(4)
$\Delta KR_{i,t}$	0.0109**	0.0152**	0.0161**	0.0162**
	(0.00547)	(0.00680)	(0.00677)	(0.00678)
$1/(1+L_{i,t})$	0.449***	0.430***	0.395***	0.408***
	(0.0120)	(0.0149)	(0.0160)	(0.0161)
$\Delta KR_{i,t} \times 1/1+L_{i,t}$	-0.0549**	-0.0716**	-0.0712**	-0.0797***
	(0.0234)	(0.0287)	(0.0288)	(0.0287)
Firm Fixed Effects	X	X	X	X
Year Fixed Effects	X			
SIC4dig×Postcode ×Year Fixed Effects		X	X	X
Firm controls (*t*)			X	
Firm controls (*t-1*)				X
Observations	901,290	900,301	865,887	869,180
N of firms	199,394	199,225	195,148	196,546
R squared	0.233	0.345	0.383	0.382

This table presents an extension of the baseline results. The dependent variable is the log difference of current liabilities by firm *i* between time *t* and *t-1*. All specifications include firm fixed-effects and either year fixed effects (column (1)) or SIC4dig*postcodearea*year fixed-effects (columns (2)-(5)). All variables are defined as in Table 4, the baseline regression.. Firm-clustered standard errors are reported in parenthesis.

Table 7: Impact of Capital Requirements on Firms controlling for Bank Characteristics

Dependent Variable: Growth Rate of Assets - Δln(TA)			
	(1)	(2)	(3)
$\Delta KR_{i,t}$	0.0138***	0.0140***	0.0144***
	(0.00468)	(0.00441)	(0.00503)
$1/1 + L_{i,t}$	0.462***	0.435***	0.463***
	(0.0114)	(0.0118)	(0.151)
$\Delta KR_{i,t} \times 1/(1 + L_{i,t})$	-0.074***	-0.075***	-0.08***
	(0.02)	(0.02)	(0.023)
Bank Size	-0.005	-0.005	-0.007*
	(0.004)	(0.004)	(0.004)
Liquidity	3.01e-4	-1e4	0.004***
	(7e-4)	(8e-4)	(0.001)
Core Funding	-7.2e-5	-5.2e-5	-1e-4
	(1.2e-4)	(1.2e-4)	(1-e4)
$\Delta Woff_t$	-5.08e-4		0.00347
	(1.9e-3)		(0.00319)
$\Delta Woff_{t+1}$		0.001	
		(0.002)	
$Size \times 1/(1 + L_{i,t})$			-0.0007
			(0.007)
$Liq \times 1/(1 + L_{i,t})$			-0.017***
			(0.005)
$CoreFunding \times 1/(1 + L_{i,t})$			0.0008
			(0.0006)
$\Delta Woff_t\, 1/(1 + L_{i,t})$			-0.0231
			(0.0145)
Firm Fixed Effects	X	X	X
Lagged Firm Controls	X	X	X
SIC4dig×Postcode ×Year Fixed Effects	X	X	X
Observations	864,751	796,353	864,751
N of firms	194,780	187,077	194,780

This table presents results for the baseline regression with the inclusion of bank characteristics. The dependent variable is the log difference of total assets by firm *i* between time *t* and *t-1*. All specifications include firm fixed-effects and the SIC4dig*postcodearea*year fixed-effects (the difference with the year fixed-effects is minimal and we do not report it). All specifications include the lagged firm variables, not shown. *Bank Size* is the natural logarithm of the bank's total assets, *Liquidity* is the ratio of the bank's liquid assets (cash+government bonds) over total assets, *Core Funding* is the ratio of customer deposits over total assets and $\Delta Woff_t$is the y-o-y change in loan writeoffs. Other variables are defined as before. Firm-clustered standard errors are reported in parenthesis.

Table 8: Impact of Capital Requirements on Firms by Bank Characteristics - 2

Dependent Variable: Growth Rate of Assets - Δln(TA)			
$\Delta KR_{i,t}$	0.0141**	0.0135***	0.0139**
	(0.00588)	(0.00516)	(0.00669)
$1/1 + L_{i,t}$	0.474***	0.458***	0.472***
	(0.0128)	(0.0114)	(0.0131)
$\Delta KR_{i,t} \times 1/(1 + L_{i,t})$	-0.074***	-0.076***	-0.085***
	(0.028)	(0.024)	(0.031)
$\Delta KR_{i,t} \times highbuf$	-0.01		-0.014
	(0.009)		(0.011)
$\Delta KR_{i,t} \times lowbuf$	0.074***		0.070***
	(0.023)		(0.023)
$\Delta KR_{i,t} \times 1/(1 + L_{i,t}) \times highbuf$	0.054		0.09*
	(0.043)		(0.05)
$\Delta KR_{i,t} \times 1/(1 + L_{i,t}) \times lowbuf$	-0.321***		-0.287***
	(0.086)		(0.092)
$\Delta KR_{i,t} \times highliq$		-0.01	-0.002
		(0.013)	(0.015)
$\Delta KR_{i,t} \times lowliq$		0.009	0.016
		(0.012)	(0.016)
$\Delta KR_{i,t} \times 1/(1 + L_{i,t}) \times highliq$		0.064	0.035
		(0.057)	(0.061)
$\Delta KR_{i,t} \times 1/(1 + L_{i,t}) \times lowliq$		-0.035	-0.073
		(0.057)	(0.071)
$\partial \Delta \ln(TA)/\partial KR$	-0.06***	-0.063***	-0.071***
$\partial \Delta \ln(TA)/\partial KR$ for highbuf	-0.016		0.004
$\partial \Delta \ln(TA)/\partial KR$ for lowbuf	-0.307***		-0.287***
$\partial \Delta \ln(TA)/\partial KR$ for highliq		-0.008	-0.037
$\partial \Delta \ln(TA)/\partial KR$ for lowliq		-0.089**	-0.128**
Firm Fixed Effects	X	X	X
Lagged Firm Controls	X	X	X
SIC4dig×Postcode ×Year Fixed Effects	X	X	X
Observations	863,521	870,168	863,521
N of firms	196,524	196,606	196,524
R squared	0.447	0.445	0.447

This table presents the results for the interaction with bank characteristics. The dependent variable is the log difference of total assets by firm i between time t and t-1. All specifications include firm fixed-effects and the SIC4dig*postcode- area*year fixed-effects (the difference with the year fixed-effects is minimal and we do not report it). All specifications include the following lagged firm variables, not shown. $highbuf$, $lowbuf$ are dummies for the distribution of the bank capital buffer at the 75th (11.7%) and 25th (6.5%) percentiles respectively, $highliq$, $lowliq$ are dummies for the distribution of the bank liquidity ratio at the 75th (2.68%) and 25th (0.87%) percentiles. The model is fully interacted and it includes the quartile dummies for buffer and liquidity on their own and in interaction with $1/(1+L_{i,t})$, but whose coefficients are not shown in the table. Other variables are defined as before. At the bottom of the table, below the coefficients, we calculate the marginal effect of KR for all groups: the group between 25th and 75th pct in the first row, above the 75th in the second and fourth rows and finally below the 25th pct in the third and fifth row . Firm-clustered standard errors are reported in parenthesis.

Table 9 – The Impact of Capital Requirements on Firms by Sector

	REAL (1)	WHOL (2)	MANUF (3)	CONS (4)	Other (5)
$\Delta KR_{i,t}$	0.0204**	0.00309	0.000795	0.019*	0.0152**
	(0.00798)	(0.00828)	(0.0106)	(0.011)	(0.00772)
$1/1+L_{i,t}$	0.575***	0.343***	0.263***	0.434***	0.427***
	(0.0190)	(0.0225)	(0.0276)	(0.028)	(0.0197)
$\Delta KR_{i,t} \times 1/(1+L_{i,t})$	-0.116***	-0.0163	-0.0314	-0.069	-0.0571*
	(0.0332)	(0.0394)	(0.0519)	(0.051)	(0.0345)
Firm Fixed Effects	X	X	X	X	X
Lagged Firm Controls	X	X	X	X	X
SIC4dig×Postcode ×Year Fixed Effects	X	X	X	X	X
Observation	289,320	159,694	129,664	115,175	292,745
N of firms	71,526	34,109	25,429	25,525	65,916
R squared	0.408	0.510	0.576	0.293	0.432

This table presents the results by firm sector. The dependent variable is the log difference of total assets by firm *i* between time *t* and *t-1*. All specifications include firm fixed-effects and the SIC4dig*postcode-area*year fixed-effects (the difference with the year fixed-effects is minimal and we do not report it). All specifications include the following lagged firm variables, not shown: $CredScore_{i,t-1}$, $Currliab/TA_{i,t-1}$ and $young, old$. Column (1) includes firms in the *Commercial Real Estate* sector only, column (2) in *Wholesale & Retail Trade,* column (3) in *Manufacturing* and column (4) in other sectors. *Firm*-clustered standard errors are reported in parenthesis.

Table 10 – The Impact of Capital Requirement by Firm Characteristic

Dependent Variable: Growth Rate of Assets - Δln(TA)				
	(1)	(2)	(3)	(4)
$\Delta KR_{i,t}$	0.0144**	0.0120	0.0159***	0.0140
	(0.00670)	(0.00796)	(0.00535)	(0.00911)
$1/1 + L_{i,t}$	0.356***	0.428***	0.417***	0.302***
	(0.0148)	(0.0145)	(0.0124)	(0.0177)
$\Delta KR_{i,t} \times 1/(1 + L_{i,t})$	-0.0776**	-0.0726**	-0.0797***	-0.0804*
	(0.0333)	(0.0364)	(0.0245)	(0.0443)
$\Delta KR_{i,t} \times 1/(1 + L_{i,t})$	0.0113			0.0142
$\times\ young\ firm$	(0.0681)			(0.0684)
$\Delta KR_{i,t} \times 1/(1 + L_{i,t})$	0.0405			0.0388
$\times\ old\ firm$	(0.0471)			(0.0476)
$\Delta KR_{i,t} \times 1/(1 + L_{i,t})$		-0.00791		-0.0177
$\times\ highrisk$		(0.0513)		(0.0547)
$\Delta KR_{i,t} \times 1/(1 + L_{i,t})$		0.0162		0.0129
$\times\ lowrisk$		(0.0479)		(0.0519)
$\Delta KR_{i,t} \times 1/(1 + L_{i,t})$			0.00824	0.00776
$\times\ highclta$			(0.0594)	(0.0641)
$\Delta KR_{i,t} \times 1/(1 + L_{i,t})$			0.0166	0.0177
$\times\ lowclta$			(0.0429)	(0.0481)
$\partial\Delta\ln(TA)/\partial KR$	-0.06**	-0.06**	-0.063***	-0.066*
$\partial\Delta\ln(TA)/\partial KR$ for old	-0.03			-0.034
$\partial\Delta\ln(TA)/\partial KR$ for young	-0.05			-0.053
$\partial\Delta\ln(TA)/\partial KR$ for highrisk		-0.063**		-0.08**
$\partial\Delta\ln(TA)/\partial KR$ for lowrisk		-0.045*		-0.0499
$\partial\Delta\ln(TA)/\partial KR$ for highclta			-0.05	-0.052
$\partial\Delta\ln(TA)/\partial KR$ for lowclta			-0.056*	-0.06
Firm Fixed Effects	X	X	X	X
Lagged Firm Controls	X	X	X	X
SIC4dig×Postcode×YearFixed Effects	X	X	X	X
Observations	870,168	870,168	868,354	868,354
N of firms	196,606	196,606	196,316	196,316
R squared	0.446	0.445	0.445	0.446

This table presents the results for the interaction with firm characteristics. The dependent variable is the log difference of total assets by firm i between time t and *t-1*. All specifications include firm fixed-effects and the SIC4dig*postcode- area*year fixed-effects (the difference with the year fixed-effects is minimal and we do not report it). All specifications include the following lagged firm variables, not shown. $highrisk, lowrisk$ are dummies for the distribution of the risk variable at the 75th and 25th percentiles respectively, $highclta, lowclta$ are dummies for the distribution of the current liabilities to total asset ratio variable at the 75th and 25th percentiles. The model is fully interacted and it includes the quartile dummies for all of these dummy variables on their own and in interaction with $1/(1 + L_{i,t})$, but whose coefficients are not shown in the table.. At the bottom of the table, below the coefficients, we calculate the marginal effect of KR for all groups: the group between 25th and 75th pct in the first row, above the 75th in the second, fourth rows and sixth row and finally below the 25th pct in the third, fifth and seventh row . Firm-clustered standard errors are reported in parenthesis.

Table 11 – The Impact of Monetary Policy

Dependent Variable: Growth Rate of Assets - Δln(TA)	(1)	(2)	(3)	(4)
$\Delta KR_{i,t}$	0.0123***	0.0164***	0.0160***	0.0361***
	(0.00437)	(0.00619)	(0.00618)	(0.00987)
$1/(1+L_{i,t})$	0.447***	0.447***	0.449***	0.492***
	(0.0114)	(0.0114)	(0.0115)	(0.0173)
$\Delta KR_{i,t} \times 1/(1+L_{i,t})$	-0.0657***	-0.0799***	-0.0793***	-0.171***
	(0.0198)	(0.0265)	(0.0265)	(0.0428)
$MPshock_t \times 1/(1+L_{i,t})$	-0.0490***	-0.0499***	-0.0488***	-0.00332
	(0.00972)	(0.00985)	(0.00988)	(0.0168)
$MPshock_t \times 1/(1+L_{i,t})$				-0.128***
$\times TIGHT_t$				(0.0381)
$\Delta KR_{i,t} \times MPshock_t$		0.00893	-0.00297	0.0364***
		(0.00918)	(0.0118)	(0.0141)
$\Delta KR_{i,t} \times MPshock_t$				-0.0773**
$\times TIGHT_t$				(0.0377)
$\Delta KR_{i,t} \times MPshock_t$			0.0173	
$\times TERMPREMIUM_t$			(0.0110)	
$\Delta KR_{i,t} \times MPshock_t \times 1/(1+L_{i,t})$		-0.0320	0.0214	-0.157**
		(0.0399)	(0.0506)	(0.0621)
$\Delta KR_{i,t} \times MPshock_t \times 1/(1+L_{i,t})$				0.348**
$\times TIGHT_t$				(0.164)
$\Delta KR_{i,t} \times MPshock_t \times 1/(1+L_{i,t})$			-0.0787	
$\times TERMPREMIUM_t$			(0.0484)	
Firm Fixed Effects	X	X	X	X
Lagged Firm Controls	X	X	X	X
SIC4dig×Postcode ×Year Effects	X	X	X	X
Observations	871,423	871,423	871,423	871,423
N of firms	196,980	196,980	196,980	196,980
R Squared	0.445	0.445	0.445	0.445

This table presents the results for the interaction with monetary policy. The dependent variable is the log difference of total assets by firm i between time t and *t-1*. All specifications include firm fixed-effects and the SIC4dig*postcode- area*year fixed-effects (the difference with the year fixed-effects is minimal and we do not report it). All specifications include the following lagged firm variables, not shown: $CredScore_{i,t-1}$, $Currliab/TA_{i,t-1}$ and $young, old$. $MPshock_t$ is a monetary policy shock constructed following a narrative approach a' la Romer and Romer (2004) for the UK economy (Cloyne (2014)). $TERMPREMIUM_t$ are the y-o-y changes in the term premium of interest rates, defined as the difference between [...]. $TIGHT_t$ is a dummy variable that takes value one if the monetary policy surprise is positive and zero otherwise. Given that $MPshock_t, TERMPREMIUM_t, TIGHT_t$ are aggregate time shocks, they are absorbed by the time fixed-effects, so their coefficients are not shown in the regression. Firm-clustered standard errors are reported in parenthesis.

Table 12 – The Impact of Multiple Banks on Capital Requirement Transmission

	Multiple	Multiple	Multiple	Single
$\Delta KR_{bank1,i,t}$	0.00385	-0.00120	-0.002	0.012**
	(0.0177)	(0.0178)	(0.019)	(0.0048)
$\Delta KR_{bank2,i,t}$	-0.0146	-0.0156	-0.007	
	(0.0122)	(0.0123)	(0.012)	
$1/(1+L_{bank1,i,t})$	0.470***	0.473***	0.470***	0.444***
	(0.0222)	(0.0234)	(0.045)	(0.013)
$1/(1+L_{bank2,i,t})$	0.154***	0.150***	0.150***	
	(0.0113)	(0.0115)	(0.011)	
$\Delta KR_{bank1,i,t}$	-0.0164	-0.0121	-0.0270	-0.055**
	(0.0478)	(0.0493)	(0.0755)	(0.022)
$\Delta KR_{bank2,i,t}$	-0.0156	-0.00974	-0.00362	
	(0.0210)	(0.0210)	(0.0215)	
Firm Fixed Effects	X	X	X	X
SIC4dig×Postcode ×Year Effects	X	X	X	X
Firm Controls		X		
Lagged Firm Controls			X	X
Observations	144,805	139,859	134,741	712,707
N of firms	27,705	27,319	26,952	165,835
R squared	0.549	0.560	0.547	0.4688

This table presents the results for firms with multiple (two) banks relationship on the effects of capital requirements. The dependent variable is the log difference of total assets by firm *i* between time t and *t-1*. All specifications include firm fixed-effects and the SIC4dig*postcode-area*year fixed-effects (the difference with the year fixed-effects is minimal and we do not report it. Firm controls are: $CredScore_{i,t}$, $Currliab/TA_{i,t}$ and $young, old$, either lagged or contemporaneous. Column (4) reports the results for firms with single banking relationship.

Table 13 – The Impact of Multiple Banks on Monetary Policy

	single	multiple	single	multiple
$1/(1+L_{i,t})$	0.402***	0.485***	0.434***	0.470***
	(0.00649)	(0.0202)	(0.0135)	(0.0332)
$MPshock_t \times 1/(1+L_{i,t})$	-0.0580***	-0.0814***	-0.0393***	-0.108***
	(0.00675)	(0.0217)	(0.0111)	(0.0337)
Firm Fixed Effects	X	X	X	X
SIC4dig×Postcode ×Year	X	X	X	X
Lagged Firm Controls			X	X
Observations	798,860	147,044	713,574	136,819
N of firms	181,277	27,824	166,135	27,089
R squared	0.467	0.544	0.469	0.543

This table presents the results for firms with single and multiple (two) banks relationship on the effects of monetary policy shocks. The dependent variable is the log difference of total assets by firm i between time t and t-1. All specifications include firm fixed-effects and the SIC4dig*postcode-area*year fixed-effects. Firm controls are: $CredScore_{i,t}$, $Currliab/TA_{i,t}$ and $young, old$

Table 14 – Impact of Capital Requirements by Sectoral Bank Dependence

Dependent Variable: Growth Rate of Assets - Δln(TA)				
	(1)	(2)	(3)	(4)
$\Delta KR_{i,t}$	0.019***	0.007**	0.03***	0.015***
	(0.007)	(0.00392)	(0.008)	(0.004)
$1/(1+L_{i,t})$	0.511***	0.429***	0.561***	0.460***
	(0.01)	(0.006)	(0.019)	(0.012)
$\Delta KR_{i,t} \times 1/(1+L_{i,t})$	-0.0703***	-0.036***	-0.134***	-0.079***
	(0.0236)	(0.014)	(0.0360)	(0.021)
$BLnonreg/BLreg$	5-e4		0.017***	
	(0.01)		(0.006)	
$\Delta KR_{i,t} \times BLnonreg/BLreg$	-4.38e-4***		-5.34e-4***	
	(1.52e-4)		(1.81e-4)	
1/(1+L)*BLnonreg	-0.003***		-0.003***	
	(2.12e-4)		(3.80e-4)	
DKR*1/(1+L)*BLnonreg/BLreg	0.001**		0.002**	
	(5.49e-4)		(8.27e-4)	
Nonbank/BLreg		-0.004		0.005*
		(0.004)		(0.003)
DKR*nonbank		-3.46e-5		-5.50e-5
		(3.25e-4)		(3.65e-5)
1/(1+L)*nonbank		-1.96e-4***		-1.89e-4**
		(5.20e-5)		(1e-5)
DKR*1/(1+L)*nonbank		4.26e-5		2.15e-4
		(1.16e-4)		(1.62e-4)
∂TA/∂DKR	-0.0507***	-0.0281**	-0.1039***	-0.0639***
∂TA/∂DKR for a one std.dev.increase in (BLnonreg) or (Nonbank)	-0.0338***	-0.0274**	-0.0721***	-0.0491***
Firm Fixed Effects	X	X	X	X
SIC4dig×Postcode ×Year Effects	X	X	X	X
Lagged Firm Controls			X	X
Observations	954,167	954,167	860,355	860,355
N of firms	209,340	209,340	193,900	193,900
R squared	0.443	0.443	0.443	0.443

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