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Felix Noth, Matias Ossandon Busch

Authors

Felix Noth
Otto von Guericke University Magdeburg,
Chair for Banking and Financial Systems, and
Halle Institute for Economic Research (IWH) –
Member of the Leibniz Association
Department of Financial Markets
E-mail: felix.noth@iwh-halle.de
Tel +49 345 7753 702

Matias Ossandon Busch
Halle Institute for Economic Research (IWH) –
Member of the Leibniz Association
Department of Financial Markets
E-mail: matias.ossandonbusch@iwh-halle.de
Tel +49 345 7753 781

Editor

Halle Institute for Economic Research (IWH) –
Member of the Leibniz Association

Address: Kleine Maerkerstrasse 8
D-06108 Halle (Saale), Germany
Postal Address: P.O. Box 11 03 61
D-06017 Halle (Saale), Germany

Tel +49 345 7753 60
Fax +49 345 7753 820

www.iwh-halle.de

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Banking Globalization, Local Lending, and Labor Market Effects: Micro-level Evidence from Brazil*

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Abstract

Little is known about how banks' domestic funding networks affect the transmission of capital flows reversals to the real economy. Our robust results show that a foreign funding shock to banks in Brazil negatively affects lending by their regional branches, especially when they are subjected to funding fragmentation. This effect triggers a sizable drop in credit and job creation at the municipal level. Our findings suggest that despite substitution possibilities across banks and firms, banks' funding networks matter to explain the distributional effects of foreign financial shocks.

Keywords: capital flows reversals, branch funding networks, bank lending; regional labor markets

JEL Classification: E24, E44, G01, G21

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1 Introduction

This paper investigates whether banks' access to intra-bank funding networks affects the link between capital flows reversals and the real economy documented in emerging countries. Global credit cycles and liquidity conditions in international financial markets have been historically linked with emerging countries' economic performance. While historical studies document a connection between current account reversals and financial crises in emerging countries since the 19th century (Reinhart and Rogoff, 2010; Bordo et al., 2010), macro-finance narratives have stressed how financial frictions such as limited pledgeability (Caballero et al., 2008) and adverse selection (Martin and Taddei, 2013) expose emerging countries to the volatility of global capital flows with negative consequences for long-term economic growth.

These explanations assign a role to domestic market structures in the banking sector in affecting the transmission of global shocks to real outcomes. Recent work by Forbes and Klein (2015) suggests for instance that macroprudential regulatory frameworks can create incentives for banks to reduce the procyclicality of credit with respect to global credit cycles. Also the the degree of financial development (Förster et al., 2014) can modify the ability of banks to moderate the transmission of capital flows' shocks. These narratives converge to the notion that banks' funding market structures can be central in explaining the widespread macroeconomic implications of capital flows shocks.

We use bank- and sub-regional-level data to test whether the real effects of a negative shock to cross-border capital flows traced via banks' foreign liabilities can be linked to the fragmentation of regional funding markets for bank branches. To explore this question, we rely on matched bank-branch balance-sheet data covering all municipal bank branches in Brazil allowing us to estimate the effect of bank-specific foreign funding shocks affecting banks' headquarters, on the credit supply of their individual branches.

Our results show that a branch connected to a parent bank that experiences a drop of foreign funding of 24% percent decreases its credit growth by 15.2% in the crisis period. Considering a standard deviation in credit growth of 25%, these estimates explain a sizable share of the sample variation in credit around the crisis. We then use variables capturing branches' access to intra-bank funding network to gauge the effect of funding fragmentation. We conjecture that branches' more distant to other related branches, or with a smaller size rank within the regions in which they operate could be impaired in their capacity to offset the shock. Alternatively, we also test whether more liquid liabilities within relevant funding networks play a role. These different measures provide compelling evidence about a lending channel that increases in the extent of branches' funding fragmentation.

We further test whether the lending channel can be offset by a substitution of credit across banks, or a substitution of real outcomes across firms. For this purpose and following similar approaches by Khwaja and Mian (2008), Schnabl (2012) or Degryse et al. (2017) we collapse the data at the municipality level by computing the average of branch-level variables weighted by their market shares. Our results show robustly that municipalities hosting a larger share of exposed branches experience a significant drop in aggregate credit and job creation (i.e., the number of formal hires) following the capital flows shock. Importantly, this result is not driven by export-intensive regions suffering from contemporaneous drops in global demand. Municipality-level measures of branches' funding fragmentation reveal that while aggregate credit is directly affected by this particular friction, job creation depends mostly on how neighbouring municipalities are themselves exposed to the shock.

Why should we expect banks' branch networks in emerging countries to play a role in the inward transmission of capital flows shocks? At first glance, regional branches could substitute away a foreign funding shock hitting their headquarter by tapping liquidity in the interbank market, by exploiting loan securitization, or by raising more retail deposits locally. On borrowers' side, firms' could switch to other lenders and compensate for the loss in funding if alternative options exist. We conjecture that despite these substitution possibilities, branches could be subjected to a form of funding market fragmentation if their access to intra-bank funding networks is limited. This would be the case, for instance, when a branch holds a rather peripheral (geographical or financial) position in its network, or when their relevant funding networks remain liquidity constrained. This is the hypothesis that we explore.

From a macroeconomic dimension, the pass-through of a funding shock exacerbated by funding fragmentation would depend crucially on firms capacity to substitute to other sources of funding. On the one side, firms exposed to liquidity shocks may switch their lenders and obtain credit from less affected banks. On the other hand, even if a credit-supply shock cannot be compensated by exposed firms, other firms operating with banks that remain shielded from the shock could step in and increase their business activity to gain market shares, so that real aggregate outcomes become negligible. Therefore, testing whether a branch lending channel fueled by funding fragmentation affects aggregate outcomes requires measures of exposure to the capital flows shock at higher aggregation levels. We perform this task by exploring whether the branch-level lending channel affects job flows at the level of Brazilian municipalities.

Exploring this research question imposes several identification challenges, ranging from the exogeneity of the funding shock to the separation between branches' credit supply and local aggregate-demand trends. We build an identification strategy using a setting in which foreign funding shocks – our observable bank-level measure of capital flows reversals – occur at the level of banks' headquarters, while lending realizes at the level of their individual branches. This approach has three main advantages. First, it enables us to separate the corporate level at which the shock takes place from the level at which outcomes are observed, thereby reducing reverse causality concerns. Second, by observing lending by each branch, we can partial out demand effects by comparing branches within municipalities. That is, we introduce municipality fixed effects in a regression with first differences of lending and foreign funding. Third, using administrative data on job flows at the municipality level, we can trace the effect of the lending channel on regional labor markets.

Similar to de Haas and van Horen (2012), Chodorow-Reich (2014) or Ongena et al. (2015), we use the collapse of Lehman Brothers in September 2008 as a cut-off point that triggers a global capital flows reversal that we can trace at the individual bank level. Hereby we exploit the regulatory data at hand that traces banks' foreign interbank liabilities throughout a sample period from January 2007 to December 2009. By focusing on the period around the collapse of Lehman Brothers we can exploit an extreme event triggered at the core of the global financial system to trace its consequences for credit and job flows within Brazilian municipalities. This approach may fail to provide unbiased estimates of the funding shock if, for instance, the drop in foreign liabilities reflects banks' own lower demand for funding. We address this concern by instrumenting the change in foreign funding by banks' pre-crisis ratio of foreign liabilities to assets, following a similar approach by Aiyar (2012).

Our paper mainly relates to a broad strand of the literature that investigates the nature and real economic consequences of cross-border capital flows reversals from emerging countries' perspective. To the best of our knowledge, our analysis provides first evidence on how funding fragmentation in regional banking markets affect the reallocation of capital in the context of a large financial shock. By linking a branch-level lending channel of foreign funding shocks with labor market outcomes, the paper contributes to understand how market structures in banking can generate distributional effects of capital flows reversals across regions.

While seminal studies on capital flows reversals focused on identifying the nature and aggregate effects of capital flows reversals (see Dornbusch et al., 1995; Calvo, 1995), other contributions have used sectoral-level data to gauge the effect of external finance dependence in real sector adjustments (see, e.g., Braun and Larraín, 2005; Gallego and Tessada, 2012). However, in these studies the specific underlying mechanisms explaining real effects remain difficult to identify. First, the lack of bank-level data linked to real sector's performance means that financial channels remain unobserved. Second, the use of sectoral data does not allow to disentangle between credit supply and demand adjustments. Therefore, both specific financial market frictions and the direction of the transmission channel – from the financial sector to the real economy – cannot be properly addressed.

Within this literature a few studies have traced the effect of liquidity shocks at the bank level on firms' performance in emerging countries. Khwaja and Mian (2008) look for instance to firm-level effects of a sudden withdrawal of U.S. dollar (USD) domestic deposits triggered by international sanctions imposed on Pakistan after the 1998 nuclear tests. Similarly, Schnabl (2012) analyzes real effects of the 1998 Russian crisis in Peru. For the period around the global financial crisis Paravisini et al. (2015) analyze how credit supply shocks affect firms' exports in Peru, while Ongena et al. (2015) use a sample of yearly matched, bank-firm-level data for Eastern Europe and Turkey to analyze adjustments to firms' outcomes that stem from banks' ex ante exposures to the crisis. Our analysis complements these previous findings in two dimensions. First, we focus on a different mechanism – banks' funding fragmentation across regions – as a possible explanation for real effects. Second, we depart from firm-level outcomes and ask whether real effects of a lending channel can also be traced at higher (sub-national) levels of aggregation. This is important as even if a lending channel exists, the substitution across banks and firms could make any macroeconomic effects negligible.

More generally, our research can also be linked to studies investigating international banking activities, the transmission of shocks between financial systems, and whether the shocks affect lending or the real sector. In particular, the notion that international banking activities can transmit financial shocks to the real economy across borders goes back to Peek and Rosengren (1997) and Peek and Rosengren (2000). Van Rijckeghem and Weder di Mauro (2001) also provides evidence of the existence of common-lender contagion effects during the Mexican, Thai, and Russian crises, and de Haas and van Lelyveld (2006) reveal that home-country economic conditions crucially determine lending by foreign-owned banks in Eastern Europe. For the case of the crisis triggered by the Lehman's collapse, studies have documented a relationship between its global spillovers and the size of bank-level foreign funding shocks (Aiyar, 2012, Noth and Ossandon Busch, 2016), or liquidity conditions in international banking networks (Cetorelli and Goldberg, 2011).

Finally, our paper contributes to a small but growing literature that examines the role of regional financial integration via bank branches for capital reallocation. Gilje et al. (2016) finds that regional bank branch networks help to integrate U.S. mortgage markets, despite of widespread asset securitization. Also for the United States, Chakraborty et al. (2018) show that banks exposed to the pre-2008 housing boom via local branches reduced commercial relative to mortgage lending in non-boom regions. Closer to our study is a paper by Bustos et al. (2017), where similar branch-level data for banks in Brazil is used to explore how regional financial integration leads to a capital reallocation effect from growing soy-producing regions to non-agricultural sectors in other regions. Our approach is different in that we focus on a different economic phenomena affecting capital reallocation across branches, namely, the characteristics of their funding networks.

2 Identification and data

2.1 Identification

We test whether the intra-bank funding network of bank branches affect the transmission of a capital flow reversal to the real economy. As a first stage of analysis, this section explores whether credit supply at the branch level reacts to a foreign funding shock affecting their headquarter bank. Finding traces of the shock in the cross-section of branches within a region would indicate that internal capital markets affect the transmission of the shock to credit supply. However, such traces could merely reflect a rather mechanical adjustment in available liquidity affecting all branches of an shock-exposed bank. Consequently, in Section 3.2 we further use variables capturing branches' funding fragmentation to test whether they affect the pass-through of the shock to local credit supply.

We estimate an econometric model linking credit growth at the municipal branch level with a foreign funding shock affecting banks' headquarters and triggered by a capital flows reversal following the collapse of Lehman Bothers in September 2008. The empirical model is represented in Equation (1):

$$\Delta\text{Credit}_{ij} = \lambda_j + \beta_1 \Delta\text{XBF}_i + \sum_{k=2}^{K} \beta_k x_{kij} + \epsilon_{ij}. \tag{1}$$

ΔCredit is the change in the log of the total amount of credit of branch i in municipality j between the pre-crisis and crisis periods. To compute this value, we take the average outstanding credit of branch i for the periods from January 2007 to August 2008 and September 2008 to December 2009. ΔCredit represents the change in the log of these monthly averages between the two periods. Our main explanatory variable is ΔXBF, which indicates the change in (log) foreign interbank liabilities of branch ij's headquarter between the same two periods. In Brazil, only banks' headquarters obtain direct funding from foreign interbank markets. Therefore, we separate the corporate level, where the shock strikes, from retail banking operations at the branch level.

To estimate Equation (1) we collapse the sample's time dimension by computing the variables' averages per period from underlying monthly data. We then compute the variables of interest ΔCredit_{ij} and ΔXBF_i as first differences. We adopt this procedure to avoid concerns about our standard errors being biased due to auto-correlation (Bertrand et al., 2004). This approach also adds simplicity to the structure and interpretation of Equation (1), because aggregated time trends and banks' unobserved, time-invariant characteristics get ruled out of the analysis by first-differentiating the data. To address the potential correlation of the error term within municipalities or across branches of the same bank, we cluster standard errors at the municipality and at the bank level.

The identification of a lending channel of the foreign funding shock must fulfill two central requirements to produce unbiased results. First, the foreign funding shock must be uncorrelated with pre-existent trends in branches' credit supply. Second, we need to exclude the possibility that the analysis is driven by demand considerations, such as by different borrower fundamentals faced by banks that experience greater drops in foreign funding. When estimating Equation (1) we address these challenges to the identification as follows.

Regarding the exogeneity assumption, we argue that the default of Lehman Brothers in September 2008 – the moment we define as the beginning of the period of capital flows reversals – is unaffected by credit supply in Brazil. This is in line with other studies that use this collapse to identify the transmission of international funding shocks (see, e.g., de Haas and van Horen, 2012, Ongena et al., 2015). Considering the structure of the empirical setting, the implied assumption is that the change in foreign funding at the headquarter level is not driven by individual events at the branch level that lead the headquarter to change its demand for foreign interbank funding. Our empirical setting, in which each branch represents a relatively small share of the banking group's assets (0.6% on average), makes it rather unlikely that a given branch matters for a banking groups' foreign funding decision. Moreover, the banks in our sample never fully halted their foreign funding practices during the crisis.

Despite these considerations, a simple OLS estimation of credit growth can be subjected to problems of endogeneity and omitted variable bias if branches from a given bank are systematically more exposed to borrowers that reduce their demand for credit in the period of analysis. The change in foreign funding may then reflect a bank's headquarter decision to cut its demand for foreign funding, as a reaction to weaker demand. We address this concern by instrumenting ΔXBF by the pre-crisis average ratio of foreign interbank liabilities to assets ($XBF/Asset$). Hereby, we build on the notion that a higher ex-ante reliance on the type of funding that was directly frozen following the events in September 2008 should plausible predict the size of the funding shock in the following period. This approach reasonably assumes that the ex-ante stock of foreign liabilities was not affected by future changes in branches' lending. The change in foreign funding is also arguable the single channel trough which the instrument can end up affecting branches' credit supply (see, for a similar approach, Aiyar, 2012).

The identification builds on the significant impact of the collapse of Lehman Brothers on Brazilian banks' access to foreign funding. Figure 1 shows the development of aggregated foreign funding (expressed in real millions of USD) of banks in Brazil, documenting a steady increase of foreign funding before September 2008 and a sharp decrease right after. The drop observed after the Lehman default in September 2008 constitutes the core of our identification strategy. Similar to Khwaja and Mian (2008), we use the varying impacts of this drought in foreign funding on banks in Brazil to investigate how the shock affects local lending through bank branches.

Another requirement for the identification is the distinction between credit demand and supply adjustments that correlate with the funding shock. To differentiate between demand and supply effects, Equation (1) includes municipality fixed effects represented by λ_j, introduced after first differentiating the data. We restrict our sample to municipalities that host at least two banks active in global interbank markets, so that λ_j holds fixed anything that is municipality-specific, such as local demand for credit. Therefore β_1 isolates the credit-supply channel linking foreign funding shocks and lending. This approach implies that Equation (1) performs a within-municipality estimation, comparing the lending outcome of two or more branches deferentially exposed to the foreign funding shock via their headquarter banks.

We select multiple headquarter- and branch-level characteristics to serve as control variables within the vector x_{kij} in Equation (1). At the branch level, we include the log of total assets and the ratios of (i) liquid assets, (ii) deposits, (iii) internal funding, and (iv) net income to total assets. These variables control for the size of branches' balance sheets, and for the features of their funding structure. The last control measures branches' profitability, which reflects also the quality of a branch's borrowers. At the headquarter level we include dummies identifying banks with a foreign owner as well as banks that are state-owned (we have taken most of the information from Claessens and Van Horen, 2014, and from banks' web-pages). Furthermore, we control for the log of total assets as a measure of size, and for the ratios of capital, liquid assets, and deposits to total assets. We also capture the riskiness of the credit portfolio at the bank level by including the ratio of non-performing loans to total credit. Table 7 provides a detailed description of all variables.

2.2 Data and descriptive statistics

We use a panel data set constructed from information on banks' balance sheets and income statements reported in call reports published by the Brazilian Central Bank. This source provides monthly data on banks' lending activity and funding structure. We integrate a data set that contains information on Brazilian banks' headquarters with the (unconsolidated) balance sheets of their individual branches located in Brazilian municipalities from the ESTBAN database. Thus, we can observe both the characteristics of the parent bank at the country level as well as the characteristics of the individual regional branches of each bank. Our sample covers the period from January 2007 to December 2009. We restrict the sample to banks with a network of municipal branches throughout the period, so that we can assess the impact of shocks on lending at individual region level. This restriction reduces the sample of 123 banks active in Brazil as of January 2007 to 100 banks.

To estimate Equation (1), we also require individual bank branches to have been active during the whole sample period. Because we observe lending at the individual municipal bank branch level, we also restrict the sample to municipalities that host at least two active branches over the sample period. This restriction is important as it enables us to saturate Equation (1) with municipal fixed effects absorbing common credit-demand trends. Furthermore, we check that the banks regularly report positive balances of foreign funding, which means we can compare banks that are similarly active in global interbank markets and that continued relying on foreign funding during the crisis. This filter underpins our interpretation of the foreign funding shocks as a supply-driven phenomenon, allowing us to focus on the intensive margin of shocks. As a final sample restriction, we drop branches with missing information for the bank traits we use as control variables, while ensuring that after this restriction, each municipality still reports the activity of at least two individual branches.

Through this screening procedure, we retain a sample of 41 banks that provide credit to 1,768 municipalities through 6,632 branches. The banks in our sample represent the largest institutions in Brazil, such that our restricted sample still represents 87.4% of the total banking assets in the country. Furthermore, the outstanding credit observed in the final sample covers 81.5% of the aggregated credit market in Brazil. The sample is less representative in the country's main financial centers though, which is to be expected, considering our focus on regional branches and retail credit. That is, banks focused solely on the investment or corporate sectors, with a larger presence in financial centers, are not represented in the sample. The average number of branches per municipality is 3.75 (or 0.12 branches per 1,000 inhabitants).

At the onset of the crisis, an average Brazilian bank held a ratio of foreign liabilities to total assets of 11.6% (10% over the total sample period). This ratio varies considerably along the foreign-ownership dimension: foreign banks report an average ratio of 18.9%, whereas domestic banks finance their balance sheet, with an average of 4.2% of foreign funding. Even though, we cannot observe the counterparts of foreign funding relationships, this latter observation suggest that foreign banks have a wider access to global interbank funding, probably funnelled via internal capital markets. The sources of this funding are concentrated in banks operating from a few countries, with the largest portion of loans being originated in the United States.[1] This link to the United States is also reflected in the fact 13 banks (out of 41) in the final sample did have, during the sample period, a correspondent institution – a bank belonging to the same banking group – operating in the United States. This information supports the idea that the foreign funding shock triggered by the Lehman collapse was funneled to the Brazilian banking system via its liquidity linkages with the United States. We empirically explore this fact in Section 5.

As many other emerging countries, the credit market in Brazil is dominated by a group of large players, including Banco do Brasil, Bradesco, Caixa Economica Federal (CEF), and Itaú-Unibanco.[2] These banks held on average 40% of total assets in the branch-level sample in the pre-crisis period. Two of these banks (Banco do Brasil and CEF) are state-owned, highlighting the predominant role of state-owned banks in Brazil. In our final sample state-owned banks report an average municipal market share of 71%. This large presence of state-owned banks matter for our analysis, as they were used during the global financial crisis to intermediate quasi-fiscal policies from the central government (see, e.g., Coleman and Feler, 2015). We therefore address the implications of these banks for our analysis in Section 5.

Foreign banks also report a relevant presence in the country, led by banks such as Santander and HSBC. Overall, foreign banks report an average municipal market share of 17.5% in the pre-crisis period. Despite their smaller size compared to state-owned banks, foreign banks are widely distributed across the country, as it can be seen in Figure 3. These banks are important in our analysis as we would expect them to have been more

[1] Data from the BIS Locational Banking Statistics (LBS) show that as of 2007, 55% of total cross-border claims to Brazil originated in the United States. Taken together, claims from the top-5 origination countries accounted for 90% of total claims to Brazil. These countries are, in order of claims' size, the United States, the United Kingdom, the Netherlands, France, and Belgium. It should be noted that these numbers pertain to claims to all economic sectors in Brazil. According to the same source, more than 70% of cross-border claims in 2007 were denominated in USD.

[2] Our sample period covers the merge of banks Itaú and Unibanco in November 2008. To avoid bias resulting from the balance-sheet adjustment driven by this merge, we assume the two banks as a single institution from January 2007 onward, that is, we merge the banks' and branches' balance sheet over the sample period.

exposed to the foreign funding shock provided their reliance to liquidity-constrained global internal capital markets (see Cetorelli and Goldberg, 2011). This conjecture is consistent with the fact that they held ex-ante a large share foreign funding. For these reasons, we carefully address the heterogeneous response of foreign banks to the foreign funding shock in Section 5.

Table 1 provides descriptive statistics for the variables in our analysis (Columns I to IV) and shows the mean values for the pre-crisis period for two groups of banks, according to whether they experienced a change of (log) foreign funding below (shock affected) or above (non-affected) the sample median (Columns V and VI). We compute normalized differences (Imbens and Wooldridge, 2009, Lambert et al., 2017) to investigate whether the differences in variables between the two groups differ significantly from each other. This difference in means is reported in Column VII.

The first three lines in Table 1 report summary statistics for the main variables of interest: the changes in log credit and log foreign funding between the two periods (ΔCredit and ΔXBF, respectively), and the pre-crisis ratio of foreign liabilities to assets ($XBF/Asset$). By construction Table 1 shows that foreign funding growth was weaker for shock-affected banks. Credit expanded in a slower fashion in the case of shock-affected banks, which report a 7 percentage points lower credit growth between the pre- and post-crisis periods. In support of the instrumental variables approach outlined above, we find that banks affected by a large drop in foreign liabilities were also the ones reporting a larger pre-crisis $XBF/Asset$ ratio.

In addition, Table 1 documents that banks affected or not by a foreign funding shock shared similar characteristics in the pre-crisis periods in most of the control variables. Some significant differences appear in the capital ratio, as well as in terms of the likelihood of being a foreign-owned bank. These statistically significant differences are marked with asterisks. Shock-affected banks tend therefore to operate ex-ante with more capital, and are more likely to be foreign-owned. For the rest of the control variables, we cannot reject the null hypothesis that the averages between the banks affected or not by the shock are equal.

A challenge to the identification strategy is the potential existence of ex ante trends in banks differently affected by foreign funding shocks. More affected banks could have been experiencing weaker credit growth in the pre-crisis period, which would prompt a bias in our estimation. The assumption of parallel trends in the pre-crisis period must be therefore addressed explicitly. In the bottom panel of Table 1, we report the results of tests of whether average pre-crisis month-on-month growth rates in credit and deposits differed significantly between the two groups of banks. This test does not indicate any statistically significant differences in pre-crisis trends between banks affected or not by the funding shock.

To confirm the validity of the parallel-trends assumption Figure 2 further provides graphical non-parametric evidence about the effect of the foreign funding shock on branches' lending. It shows the change in aggregated log outstanding credit for groups of banks reporting a measure of ΔXBF above and below the sample median. Credit growth is computed as proportional to outstanding credit as of September 2008. Figure 2 shows no diverging pre-trends in lending between these two groups of banks. The figure also shows a drop in credit growth after the crisis' outbreak by affected banks, in line with our hypothesis.

3 Branch-level lending channel

3.1 Benchmark results

Table 2 reports the benchmark results from estimating Equation (1) using the $XBF/Asset$ ratio as an instrument of ΔXBF around the crisis period. Column I reports the main specification in which we use the log change in total outstanding credit for each branch to compute ΔCredit. The positive and statistically significant coefficient can be read as follows: a 1% decrease foreign funding after the crisis led to a significant reduction in the growth rate of lending of about 0.62%. Considering that shock-affected banks (i.e., those with ΔXBF below the median) experienced an average drop in foreign funding of 24%, the foreign funding shock leads to a drop of roughly 15% in credit by the average branch ($20\times0.63=15.2$). Compared this number with a standard deviation of 25% in ΔCredit, the model explain a sizable portion of credit growth variation in the sample.

Consider now the difference between the average values of ΔXBF for affected and non-affected banks. Our estimates imply that, on average, credit growth was 51% (82×0.63) lower for affected banks as a consequence of the shock. If an average non-affected bank would had realized the foreign funding growth rate of an average affected bank, ΔCredit (13% on average) would have been more than three times lower (13%-51%=-38%). This example illustrates the extent of the effect of the shock on local credit supply. The documentation of this bank lending channel for the Brazilian financial system mirrors the findings of other studies that analyze how funding shocks affect banks' lending behavior (e.g., Khwaja and Mian, 2008, Schnabl, 2012), in particular in the context of the global financial crisis (e.g., Ongena et al., 2015, Chodorow-Reich, 2014). Our findings highlight how intra-bank liquidity channels lead to branch-level credit-supply adjustments when funding shocks occur.

Columns II to V in Table 2 replicate the analysis by computing ΔCredit for different subsets of credit segments: commercial loans (ΔC&I), consumer loans (ΔCons), mortgages (ΔMort), and leasing (ΔLeas). The results show that both commercial and consumer loans are sensitive to ΔXBF (Columns II and III), whereas for mortgage loans and leasing the null hypothesis cannot be rejected. This finding may reflect the importance of collateral in retail credit markets, especially during a global financial crisis (Ongena et al., 2015). Whereas mortgages and leasing products can insure banks against repayment delinquency, the other categories of credit do not necessarily provide this function.[3]

This credit-segment view is important as it addresses concerns that the fixed effects specification in Equation (1) may fail to absorb demand-driven variations in credit. This would be the case if, for instance, branches operate with focuses in different credit segments, implying that municipality fixed effects cannot absorb rather branch-specific credit demand dynamics. The fact that the results can be confirmed for commercial and consumer loans indicates that even within specific credit markets in each municipality, credit volumes decrease by more for more affected branches.

The coefficients for the control variables in Table 2 report economically meaningful results. In line with previous findings by Coleman and Feler (2015), we find credit supply to be positively correlated with government ownership of banks, evidencing a potential offsetting effect of government-owned banks interventions in local credit markets. Also a higher exposure to credit risk is associated with a weaker credit growth. At the branch level, a larger balance sheet's size and a higher profitability tend to shield credit supply from the shock. Finally, branches whose headquarters report larger liquidity ratios reduce credit by more, what may be related to a liquidity hoarding effect, as documented for the global financial crisis in the United States by Cornett et al. (2011) and Berrospide (2013).

3.2 Branch lending and funding networks

The results in the previous section document a branch-level lending channel in Brazil arising from a foreign funding shock triggered by the collapse of Lehman Brothers in 2008. However, the transmission of funding shocks from parent banks to branches' credit supply is not obvious a priori. First, the crisis could have generated heterogeneous demand shocks across branches and regions, affecting the volumes of credit granted. While we

[3]We expect this to be a relevant factor in Brazil considering the theoretical and empirical evidence on the importance of collateral for credit markets in emerging countries compared to developed countries (e.g., Fostel and Geanakoplos, 2008, Menkhoff et al., 2006). In support of this interpretation, we note that unlike the United States, Brazil did not experience a housing bubble before or during the crisis.

are confident that the model in Equation (1) reasonable captures demand trends within regions, Section 5 reports several robustness tests that confirm the validity of the reported estimates. Second and most important for our research question, affected branches may have had access to alternative funding sources so that the shocks could have led to a simple shift out of internal funding and towards local deposits.

In this section, we extend our baseline model to explore whether the capacity of branches to access alternative funding within its intra-bank funding network explains the extensive margin of the credit supply adjustment. Following previous theoretical literature, we construct measures of funding market fragmentation capturing the geographical and financial position of branches within their relevant intra-bank funding networks. If, as we expect, a branch obtains much of this funding from the same banking conglomerate to which it belongs to, then branches distant to intra-bank liquidity pools should experience a larger adjustment in credit supply.

The hypothesis that branches' funding market fragmentation affects the transmission of the foreign funding shock could be, at first glance, easily testable a priori. For example, one could compute measures of internal capital markets' dependence and check whether the identified effect increases in the distribution of branches' pre-crisis internal funding ratio. However, this approach would be subjected to two important methodological drawbacks. First, the ratio of internal funding to assets could reflect branches' business models characteristics or changes in the bank-branch relationship occurring in the pre-crisis period that could be correlated with the shock's transmission. Second, internal funding per se would fail to capture the characteristics of the funding network relevant for each branch, including the closeness to related parties and the (intra-bank) liquidity available in its geographical vicinity.

We therefore follow an approach based on measures of funding market fragmentation that, departing from branches' own balance sheet structures, reflect their access to alternative funding via intra-bank networks. First, we use geographical coordinates obtained from the Brazilian Institute of Geography and Statistics (IBGE) to compute the log of the distance between a branch and the closest branch of the same bank. This measure captures information and financial linkages that we would expect to be tighter between closer branches, allowing for a swifter reallocation of capital. This variable builds on the notion that local liquidity pools can be seen as a first line of defence when large-scale funding shocks occur (see, i.e., Allen and Gale, 2000).

Second, we compute the log of branches' size rank within each federal state. This variable captures the hierarchical structure of branches at the sub-national level. If branches' funding network matters for the reallocation of capital, we would expect branches in a higher hierarchical position to be able to manage and exploit liquidity channels from a consolidated perspective, benefiting from a more central position within their regional funding networks. On the contrary, smaller branches are likely to be subjected to funding constraints due to their peripheral position in the network.

While these latter variables measure geographical characteristics of funding networks (i.e., closeness to related entities and centrality in the regional network hierarchy), financial fragmentation could also be reflected in the available liquidity in these networks. Even if a branch is geographically distant from liquidity centers, it may benefit from the fact that a few related branches in its immediate vicinity dispose of large liquidity pools that can be shared when liquidity constraints arise. We therefore compute variables capturing the extent of available liquidity in a branches' funding network. First, for each branch i, j we compute the ratio of aggregate deposits (sight + saving) to assets of all branches different than i, j within the same bank i. We use this broad measure of available liquidity as a sense check to test whether the liquidity in intra-bank funding networks can be related to our benchmark results. We then compute the average ratio of deposits to assets of all branches different than i, j of the same bank within a given micro-region.[4]

We use the variables described above to look for traces of financial fragmentation in the relevant funding network of each branch. We therefore select these variables to capture relevant aspects of banking networks and financial contagion previously identified in the theoretical literature. The distance variable can be related to the notion of complete versus incomplete networks in financial contagion introduced by Allen and Gale (2000). In our empirical setting, the network's "regions" are represented by branches, whereas the overlapping claims held by regions in Allen and Gale (2000) translate in our case to a "common lender effect" (Kaminsky and Reinhart, 2000) due to the branch-headquarter relationships. Our hypothesis on the role of financial fragmentation follows the argument by Allen and Gale (2000) that banks in incomplete (i.e., financially disconnected) networks are more exposed to liquidity risk as the network fails to prevent risk-sharing mechanisms when a shock occurs.

Beyond distance, the other variables capture additional angles of branches' funding network characteristics. First, the regional size rank provides a complementary measure of the centrality of a given branch within its relevant network. This variable also adds a core-periphery and hierarchical definition to a branch's position vis-à-vis other related branches (see, e.g., Borgatti and Everett 1999). Second, the variables measuring the availability of

[4]Micro-regions were, at the time of the sample period, legally defined geographic areas grouping a small number of municipalities. These areas were defined mostly for statistical purposes according to their economic, geographic, and cultural characteristics by the IBGE. In the sample, each micro-region consist, on average, of 3.4 municipalities. In Table A.21 in the Online Appendix we report descriptive statistics for the variables measuring branches' financial fragmentation.

liquid liabilities (i.e., deposits) in a branches' funding network add a financial perspective to the definition of fragmentation. Hereby we build on the notion of harmonic distance introduced by Acemoglu et al. (2015) in which the distances in a network are weighted by the extent of bilateral exposures. In our setting what matters are not bilateral claims per se (as the shock propagates unilaterally from the parent bank) but rather the available (short-term) liquidity in the relevant funding network. We expect more liquid nodes within a network to provide access to wider buffers to compensate for shocks, as in Allen and Gale (2000). The stabilizing role of insured deposits for banks' funding risk has been previously discussed by Acharya and Naqvi (2012) and Khan et al. (2017).

The results of this analysis are reported in Table 3. For comparison, we replicate in Column I the benchmark results reported in Table 2. The subsequent columns report regressions in which we add to Equation (1) an interaction term between ΔXBF and the financial fragmentation variables mentioned above. Columns II to V use as an interaction term the (log) distance to the closest branch (Distance), the (log) size rank of branch i, j within its bank and federal state (Size rank), the pre-crisis ratio of aggregate deposits to assets by all branches of bank i excluding branch i, j (Bank deposit), and the average deposit ratio of all other branches of bank i within a branch's micro-region (Regional deposit). These tests provide evidence consistent with the hypothesis that funding fragmentation leads to a stronger pass-through of the funding shock to lending.

We interpret this findings as follows. First, branches located in the periphery of their relevant funding network (i.e., those more financially fragmented) reduce credit supply by more than other similarly affected banks. This is true for branches with a larger distance from their closest related entity or with a smaller size rank and therefore lower hierarchical position. Second, branches related to funding networks with a wider availability of local deposits (in opposition to, i.e., internal funding) either at the country or at the micro-region level report a smaller effect on credit supply (negative sign on interaction terms in Columns IV and V).

To gauge the effect of financial fragmentation consider a bank with an average distance to its closest related branch. The estimated effect of ΔXBF on that branch (at the average of ΔXBF=0.24 for affected banks) is a 8.8% lower growth rate in credit when compared to a non-affected branch (3.24×24×0.113=8.8). Alternatively, a similar branch with a *Distance* value 1 s.d. above the mean reports an estimated effect on credit growth of 11% (4.04×24×0.113=11). Therefore, increasing the distance to the closest related branch in 1 standard deviation leads on an increase in the estimated effect of 25% (2.2/8.8=0.25). We interpret this results as evidence of a sizable effect of financial fragmentation on the pass-through of the foreign funding shock.

The results from Table 3 point out that beyond branches' own funding structure, their position within their local funding network matters for the transmission of a large capital flows reversal to credit supply. This finding is consistent by previous evidence on the sensitivity of bank branches to the performance of their banking conglomerate (e.g., Houston and James, 1998, Boutina et al., 2013, Giroud and Mueller, 2017, Bustos et al., 2017). Our results add to this literature by suggesting that beyond own branches' liquidity constraints, funding network characteristics are important vectors for the transmission of financial shocks.

4 Real effects of the lending channel

The results from Section 3 suggest that large capital flows reversals can be transmitted to the real economy via branches' credit supply. However, providing a full picture supporting this narrative requires tracing and quantifying the potential real economic effect of the documented lending channel. This is important for two reasons. First, borrowers could compensate for a shortfall in credit from one affected bank by tapping funding in another, less affected bank. Also less affected firms could exploit opportunities and increase their outcome (i.e., investment, hires). This would make any real economic implications negligible. Second, this analysis can be used to test whether regions hosting branches with higher funding fragmentation react to the credit-supply shock by more.

In this section we test if and how real outcomes at the municipality level were affected by a shock to the foreign funding position of banks active in those regions. For this purpose we include in the sample all bank branches in the municipalities from our benchmark sample excluded in the sampling procedure. In doing so, we allow for the possibility that borrowers may offset the lending restriction imposed by shock-affected banks by accessing credit in other banks. For this analysis, we aggregate the data at the municipality level by weighting bank and branch variables by the share of each bank in a municipality's credit market.[5] With this data set, we run the following regression:

$$\begin{aligned}\Delta \text{Log outcome}_j = {} & \alpha_0 + \alpha_1 \Delta \text{XBF}_j \\ & + \sum_{k=2}^{K} \alpha_k x_j + \gamma_{micro} + \epsilon_j,\end{aligned} \tag{2}$$

[5]If a bank has missing data related to its foreign funding position, we impose an assumption that the bank experienced a value of ΔXBF of 0. We need to retain banks that do not report regularly active positions of foreign liabilities in the sample to obtain conservative estimates of the borrowing channel of financial contagion. If we instead consider only the 41 banks from the baseline sample, we would only allow customers to switch off their funding sources across those banks. The final sample including all banks features 100 banks and 11,134 bank branches in the same 1,768 municipalities of Section 3.

where outcome refers to real outcome variables at the municipality level j. In different regressions, we estimate the change in log aggregate credit by all branches in municipality j ($\Delta Mcred$), the change in the log of job creation (ΔJC), and the change in the log of job creation per 1,000 inhabitants (ΔWJC). These variables are computed as changes between the pre-crisis and crisis periods. For this purpose we compute the average number of each value for the respective period from the underlying monthly data at hand. The main variable of interest remains ΔXBF, which is computed as a market-share weighted average of the value of ΔXBF reported by all branches within municipality j.

By aggregating the data at the municipal level the empirical strategy to absorb credit-demand shocks via municipality fixed effects cannot be implemented. However, we still can exploit the sub-regional structure in the data to run Equation (2) with micro-region fixed effects (γ_{micro}). To the extent that the rather small number of municipalities within a micro-region shared common economic, political, and cultural characteristics, we expect these fixed effects to capture aggregate-demand factors that could explain $\Delta \text{Log outcome}_j$ via alternative channels. By following this procedure we estimate Equation (2) within micro-regions, comparing municipalities with different ΔXBF but sharing similar other characteristics.[6]

Three important features of Equation (2) should be noted. First, by introducing the term γ_{micro} after first-differentiating the data the constant α_0 captures all secular time trends in the economy at the micro-region level. Second, even in this setting it may be the case that within a micro-region, municipal outcomes are correlated with ΔXBF. We therefore estimate Equation (2) replicating the instrumental variables' approach used in Equation (1), that is, we instrument ΔXBF by the $XBF/Asset$ ratio computed as the market-share weighted average of the pre-crisis $XBF/Asset$ ratio reported by all branches in j. Third, we use the vector of control variables x_j to capture municipalities' relevant pre-existent characteristics. Due to the aggregation procedure, we control for instance for the market share of foreign or state-owned banks, for the available liquid assets in the local banking system (that can be used to compensate for asset losses), and for the size of the deposit base that can be used by firms and households. Additionally, we use the log of 2007 GDP to absorb the effect of regional size in ΔLog outcome. When estimating

[6]Micro-region fixed effects can capture, for instance, whether municipalities are located in economically important regions (i.e., important for the production of commodities), or whether certain economic sectors such as mining or agriculture denominate. Moreover, we expect γ_{micro} also to capture whether a municipality is located in a rather export-intensive region. This latter fact matters because the crisis could have arguably triggered a negative credit-demand shock in regions directly exposed via the export industry to the United States. In Section 5.2 we implement several robustness tests to address this critique. In Section 5.1 we show an OLS version of Equation (1) indicating a fairly stable size of the coefficient for ΔXBF in the branch-level regressions (0.62 in the fixed effects versus 0.74 in the OLS model). This fact should further reduce concerns of a significant demand-driven bias in municipality-level analysis.

Equation (2) we cluster standard errors at the micro-region level.

As a first step in the analysis, we test whether borrowers were able to compensate for the shock by switching their funding sources. We test this hypothesis by estimating the effect of ΔXBF on the change in log aggregated outstanding credit at the municipal level (ΔMcred). The results in Columns I (OLS) and II (with γ_{micro}) of Table 4 provide evidence against the hypothesis of cross-branches substitution: within a given micro-region, municipalities with a larger exposure to the funding shock (i.e., lower ΔXBF) experience weaker credit growth ex-post, a result that is statistically significant. Importantly, the estimated coefficient remains statistically significant and stable in size in regressions with and without FE. We thus have initial evidence against the substitution of credit, opening a path for further consequences in local economies.[7]

4.1 Lending channel and labor market outcomes

Next, we explicitly investigate whether the lending channel can be associated with real economic adjustments during the crisis period. For this purpose we compute the the change in the log of gross and population-weighted job creation (ΔJC and ΔWJC, respectively) using administrative data from the Brazilian Ministry of Labor, which reports these statistics under the General Survey of Employed and Unemployed (Cadastro-Geral de Empregados e Desempregados). We define WJC as the number of hires per 1,000 inhabitants using year-end population statistics from the Brazilian Institute of Geography and Statistics.

The time series of job creation reveal a large effect of the crisis on the Brazilian labor market (see Figure A.2 in the Online Appendix for details). On aggregate, the Brazilian economy was creating an increasing number of job contracts in the crisis' run-up (2007-2008), reaching a maximum of 12 jobs per 1,000 inhabitants per month in September 2008. After that, the data evidences a major drop in hires, which fell to around 9 contracts per 1,000 inhabitants in January and February 2009. On average, job creation fell by 20% between the pre- and post-crisis periods. Hires also fell net of layoffs: while net job creation was positive throughout the pre-crisis period, it remained in negative values since the crisis' outbreak and almost entirely throughout 2009. This effect was large from an historical perspective, as longer times-series corroborate. In fact, the drop in job creation around 2008-2009 was the largest since 2000.[8] These facts suggest that the global financial crisis had a sizable effect on the Brazilian labor market.

[7] Note that regressions using micro-region fixed effects reduce the number of municipalities from 1,768 to 1,640 since only micro-regions with at least two municipalities enter these regressions.

[8] The fact that job creation converge to pre-crisis levels by 2010 highlights the anomaly represented by the crisis, a feature that underpins the identification strategy.

To gauge the effect of municipality-level ΔXBF on local labor markets we estimate Equation (2) using job creation growth rates as dependent variables. To assess the implication of micro-region fixed effects, we report for each dependent variable regressions including and not the γ_{micro} term. The results are reported in Table 4. Columns I and II report the aforementioned effect on aggregated credit growth. Then, Columns III and IV display regressions using the long change in average job creation (ΔJC) as the dependent variable. The positive and statistically significant coefficient indicates an empirical link between the drop in hires during the crisis and the extent of the shock ΔXBF. This result holds and remains fairly stable in size irrespective of the inclusion of micro-region fixed effects.

Columns V and VI in Table 4 report the results when defining the dependent variable as the population-weighted growth rate in job creation (ΔWJC). These regression provide more economically meaningful results, since the change in hires is weighted by municipalities' size. These regressions confirm our previous finding, with the coefficient for ΔXBF reporting a positive and statistically significant value, similar in size to the unweighted regressions. Taking the regression in Column VI as our benchmark result, we conclude that a 1% decrease in ΔXBF is associated with a 0.32% decrease in job creation by 1,000 inhabitants on average. This result implies that a shock equal to one standard deviation in ΔXBF (i.e., 19%) would lead to a drop in job creation of 6.08% (19×0.32=6.08). This effect represents around 33% of one standard deviation in ΔWJC (with one standard deviation equal to 18.7%), and explains therefore a sizable share of the within-sample variation in job creation.

The branch lending channel identified in Section 3.1 is thus by no means innocuous. When borrowers fail to access alternative funding sources to substitute for their reliance on affected banks, the lending channel can have significant effects on the real economy. Our findings are in line with previous studies linking bank-level funding shocks with real outcomes at the firm level (see, e.g., Schnabl, 2012, Paravisini et al., 2015, Ongena et al., 2015). We add to this literature by showing that real effects can also occur at higher levels of economic aggregation, despite substitution possibilities. The regional-level view in our analysis also captures possible negative spillovers across firms (see a similar discussion in Huber, 2018), corroborating evidence by Gabaix (2011) and Amiti and Weinstein (2018) on the macro implications of granular (i.e., bank-specific) financial shocks.

4.2 Municipality-level effects and funding networks

A natural question is whether branches' funding fragmentation can also directly explain the transmission of the credit-supply shock to municipality-level outcomes. We therefore implement a set of regressions in which we adjust Equation (2) by adding an interaction term between ΔXBF and the proxies for branches' funding market fragmentation used in the branch-level analysis. These variables are computed at the municipality level as market-share weighted pre-crisis averages.

Since the substitution of the shock by firms may depend on variables different than their own branches' financial fragmentation, we include in this exercise a further regression in which ΔXBF is interacted with the average ΔXBF of all other municipalities different than j within the same micro-region. We follow this approach to allow Equation (2) to test – in a horse-race fashion – whether the relevant economic force affecting the possible substitution of the shock is either driven by branches' financial fragmentation or, alternatively, by the size of ΔXBF in neighbouring regions.[9]

Ex-ante it is not straightforward to predict the results of adding this interaction term to estimate aggregate outcomes. On the one hand, financial fragmentation may for instance affect job creation exclusively via its effect on credit supply. Similar to the exclusion restriction in an instrumental variables' model, this hypothesis would imply that the branch-level results for financial fragmentation from Section 3.2 cannot be replicated when testing the substitution across banks and firms in Equation (2). On the other hand, financial fragmentation could affect both branches' credit supply and the substitution of credit or hires. This latter case would hold only if relevant substitution channels occur exclusively within municipalities and not across them: if some firms can access funding in other regions, then the funding fragmentation within municipality j in which those firms operate should not have a direct effect on macro outcomes (beyond its indirect effect via credit).

The results from this exercise are reported in Tables 5 and 6 for ΔMcred and ΔWJC, respectively. Columns I to IV report the results of regressions with the financial fragmentation proxies used in the branch-level analysis. In Column V the interaction term is the average ΔXBF in all municipalities different than j within the respective micro-region (ΔXBF-reg). When comparing Tables 5 and 6 an interesting pattern emerges. While the interaction terms in Columns I to IV render statistically significant coefficients in line with the branch-level results for ΔMcred, this is not the case for ΔWJC. However, when the interaction term is ΔXBF-reg (Column V) the conclusion is the opposite: while this variable seems not to affect aggregate credit, it does have an effect on ΔWJC with the

[9] It should be noted that adding the interaction term between ΔXBF and fragmentation proxies also provides a robustness test for the results from Section 4.1. If, for instance, the effect of ΔXBF is fully absorbed by the variable *Distance*, it would mean that the substitution of credit and labor does take place, as long as the average distance to related branches in a given municipality is low. This result would challenge our conclusion above against the existence of substitution of labor or credit in the affected municipalities.

expected sign, i.e., the effect increases when neighbouring municipalities are more exposed to the shock.

These results highlight an important conclusion of our analysis. When analyzing firms' capacity to substitute away the shock by taping liquidity in other branches, the financial fragmentation variables from Section 3.2 matter. Therefore, it is the possibility of substitution within municipality j the one that affects aggregate trends of credit volumes. However, the results also suggest that the economic forces behind the effect of the lending channel on ΔWJC are different. Here, the non-linear effect over ΔXBF-reg implies that when neighbouring municipalities are less exposed to the shock, firms report a lower drop in hires. This latter conclusion has two interpretations. First, branches' financial fragmentation matters for the real economy only via its effect on credit supply. Second, when branches in neighbouring regions are less exposed to the capital-flows reversal firms do partially substitute the shock and the effect on job creation decreases.[10]

5 Robustness tests

In this section we describe robustness tests aimed at testing the validity and stability of our results under alternative specifications. We distinguish between tests addressing concerns with the identification of the lending channel (Section 5.1), and tests on the municipality-level credit and labor market results (Section 5.2). We report these tests in the Online Appendix.

5.1 Robustness tests: branch-level lending channel

We acknowledge that the instrumental variables' approach in Equation (1) can be subjected to usual critiques regarding both the relevance and the exogeneity assumptions implied by the instrumental variables' estimation. To address these concerns we divide the instrumental variables' approach into different stages that illustrate the validity of these latter assumptions. The results are reported in Table A.1. First, we estimate the first stage of Equation (1), with ΔXBF as the dependent variable and the $XBF/Asset$ ratio as the main explanatory variable. Consistent with our previous discussion, we find

[10]One limitation of the municipality-level data is that we cannot directly explore which firms are leading this substitution. One possibility is that the substitution across municipalities reflects that less-affected firms (i.e., firms whose branches are less exposed to the shock) hire more workers to exploit business opportunities when other municipalities in the micro-region remain financially resilient to the shock. Alternatively, affected firms may exploit the fact that a neighbouring region is less financially constrained to obtain alternative funding there. This dynamic would reinforce the drop in aggregate credit in municipality j and simultaneously lead to a stronger job creation.

that a larger pre-crisis $XBF/Asset$ ratio predicts a lower growth rate of foreign funding after September 2008.

We then perform a test in which we estimate Equation (1) for the subsample of banks whose headquarters report a pre-crisis $XBF/Asset$ ratio below the 25th percentile of the banks' sample distribution. Following Angrist et al. (2010), we expect these banks to be "never-takers", in the sense that the instrument $XBF/Asset$ should not predict ΔXBF for those banks. The results confirm this hypothesis. We further estimate a reduced-form model of Equation (1) in which we regress ΔCredit directly on $XBF/Asset$ for the subsample of low-$XBF/Asset$ banks, finding no statistically significant results. These tests suggest that the instrument is only informative about the size of shocks and effectively identifies a lending channel for banks with a relatively large $XBF/Asset$ ratio. Moreover, these results confirm that our estimation does not arise from a rather spurious correlation between ΔXBF and $XBF/Asset$.

We then test the stability of our results when changing the characteristics of the econometric model or when excluding relevant observations. This sensibility analysis supports the validity of our main findings under different specifications. First, we consider a version of Equation (1) including only bank-level controls in the vector x_{kij}. Second, we estimate the model by adding branch controls but excluding municipality FE. Third, we lag the vector of control variables in one period, so that control variables enter the model as pre-crisis averages. Third, we compute ΔCredit as the change in average credit volumes between the two periods scaled by a branch's pre-crisis total assets, weighting the dependent variable by branches' size (i.e., log assets).

We also estimate Equation (1) with alternative clustering for the standard errors at either the bank or the municipality level. Finally, we estimate Equation (1) by excluding relevant observations that could be arguably biasing our results. Consequently we exclude (i) all state-owned banks, (ii) the main federal states of Brazil (i.e., Rio de Janeiro, Sao Paulo, and Minas Gerais), and (iii) all the capital municipalities of the federal states. These latter tests confirm that neither state-owned banks nor specific dynamics in the capital regions or cities that may be arguably correlated with ΔXBF influence the results. These results are reported in Tables A.2 and A.3.

One limitation of Equation (1) is that it cannot be used to explicitly test whether the lending channel reflects other underlying characteristics of banks, branches, or regions that could potentially intensify or moderate the extent of the credit supply effect. For instance, if the funding shock effectively arises as a consequence of the events triggered in the United States, we would expect banks with tighter U.S. institutional linkages to be more affected. To explore possible nonlinearities we therefore extend Equation (1) by including an interaction term between ΔXBF and other variables that could be arguably playing a role in the credit supply adjustment. In Table A.4 we report regressions using the following interaction terms: a dummy equal to 1 for private banks; a dummy equal to 1 for domestic banks; a dummy equal to 1 for non-U.S.-connected banks (both foreign and domestic); a dummy equal to 1 for banks with a pre-crisis average ratio of foreign investments to assets above the median; the average ratio of branches' internal funding to assets ratio; the average log of banks' and branches' assets; the average log of aggregate assets per municipality; and the log of municipal population. These variables are measured in the pre-crisis period (end of 2007 for population).

Our benchmark results are robust to including these interactions terms, which also provide further evidence on relevant underlying mechanisms. First, the effect is driven by private banks, confirming (i) that state-owned banks were shielded by the effect (see Coleman and Feler, 2015), but (ii) that the effect is not explained by low-$XBF/Asset$ state-owned banks. Second, the results are stronger for foreign banks, and in particular for banks with U.S.-ties via the presence of related institutions (i.e., headquarters, affiliates, or other U.S.-based branches or subsidiaries of the same bank) in the United States. Foreign investments exposures, if anything, reduce the size of the effect, confirming that the results are not driven by asset losses abroad. We also find that branches with a higher ex-ante ratio of internal funding show a stronger sensibility to ΔXBF, in line with the notion that internal capital markets are transmitting the funding shock to branches. We do not find evidence of neither banks' or branches' size affecting the documented lending channel. However, branches operating in larger municipalities (i.e., larger aggregate bank assets or population) do react to a larger extent.

Our identification strategy relies on the assumption that branches within a municipality are effectively competing in a similar pool of borrowers to capture credit demand trends. In fact, one may argue that two branches operate in a rather oligopolistic credit market in which certain portions of the potential borrowers are divided ex-ante between branches. In that case, credit demand trends would be at least partially branch-specific, weakening the credit-demand control represented by the term λ_j in Equation (1). While the credit-segment specific regressions in Table 2 should reduce this type of concerns, we acknowledge that exploring this competition assumption requires further analysis. Therefore, we implement a test in which we exclude municipalities in which we expect banking competition to be relatively low (i.e., below the median) of different competition measures. In this test we exclude the following municipalities: small municipalities as measured by the log of aggregate bank assets or the log of the number of branches; municipalities with a low average market share of all its branches; and municipalities with a high Herfindahl-Index measured from branches' market shares (all variables are measured as pre-crisis averages). These tests confirm that our findings can be replicated even within the subset of municipalities with expected high competition (see Table A.5).

Equation (1) would also fail to absorb credit demand trends if (i) the effect is stronger in regions where the original (U.S.-originated) shock generates a direct drop in aggregate demand and (ii) if simultaneously most affected branches are mainly exposed to the sector experiencing the negative demand shock. This is the case if affected branches operate mostly in export-intensive regions and are also disproportionately exposed to exporting firms. Then, our results could be explained by a (branch-specific) drop in demand that reduces credit supply for reasons beyond the foreign funding shock itself.

To address this critique we estimate Equation (1) by excluding municipalities that could be arguably the ones with the largest exposure to the drop in exports triggered by the crisis. We consider three measures of exports' exposure: the pre-crisis average ratio of exports to GDP, and the log change in average total exports and average exports to the United States around the crisis. We compute these variables using public administrative data on exports at the municipal level provided by the Brazilian Ministry of Commerce. We then exclude municipalities above the 90th percentile in the exports-to-GDP ratio or below the 10th percentile in the growth rate in exports. These results, reported in Table A.6, show that the main findings are robust to excluding export-intensive municipalities.[11]

A further critique to the empirical setting is that several banks report between the two periods an increase in foreign funding (i.e., a positive value of ΔXBF). This fact is reflected in a positive average value of ΔXBF when considering the full sample (see Table 1). This fact is not surprising as we would expect the aggregate drop in foreign funding observed in Figure 1 to reflect the loss in funding by certain banks with a stronger ex-ante exposure to the crisis, for instance via their institutional ties to the United States. However, a large average value of ΔXBF could be problematic if it lead our results to be explained by high-ΔXBF banks increasing their credit supply relative to other banks. To address this concern we follow two alternative approaches. First, we truncate ΔXBF by replacing all positive values by 0. Second, we drop the top-10 banks in ΔXBF, bringing the average ΔXBF to a value of -0.03. These regressions, reported in Table A.8, confirm that our results are not affected by banks experiencing an increase in ΔXBF. We also reconsider the role of M&As and banks' institutional characteristics for our analysis. In particular, we tests whether our results are affected by (i) the merge of banks Itaú and Unibanco in November 2008, by (ii) the presence of banks from the same banking

[11]This tests acknowledges the skewness of the distribution of exports across municipalities. For example, the average export-to-GDP ratio for municipalities above the 90th percentile is 39.8%. For the subsample below that threshold the average ratio is of 4.3%. As an alternative approach we split the sample according to the median of the export-exposure measures and report the results for the subsample of municipalities with a large versus low value for each variable, confirming the robustness of the results (see Table A.7).

conglomerate, or by (iii) the inclusion of credit unions as a special type of bank ownership. Regarding the first concern (i), the empirical setting assumes Itaú and Unibanco to be merged from the beginning of the sample period by adding up their bank- and branch-level balance sheets from 2007 onward. To confirm that this event does not affect our results we estimate Equation (1) by dropping Itaú-Unibanco from the sample, finding similar results. The second concern (ii) affects two banks in our working sample that belong to the same bank, namely Itaú Unibanco and the investment branch of Itaú called Itaú BBA. Our results remain unaltered when considering these two banks as a single entity. Finally, our sample includes one credit union (Banco Sicredi) which, even though small in size, could still affect our analysis. However, our results can be replicated by excluding that bank. These results are reported in Table A.9.

5.2 Robustness tests: municipality-level effects

To assess the robustness of the municipality-level effects we first provide a sense check for the validity of the relevance and exogeneity assumptions behind the instrumental variables' model in Equation (2). For this purpose we report the first stage of the model, in which we regress ΔXBF on $XBF/Asset$ when these variables are computed as market share weighted averages at the municipal level. This regression, reported in Table A.10, shows that a high $XBF/Asset$ ratio affects negatively the value of ΔXBF. We then estimate the first stage and the reduced-form model of Equation (2) for the subsample of municipalities with a pre-crisis $XBF/Asset$ ratio below the 25th percentile of the municipalities' sample distribution. The reduced form model directly regresses ΔMcred and ΔWJC on $XBF/Asset$. While the first stage still renders negative and statistically significant results for this subsample, the lack of significance in the reduced form regressions suggest that the outcome variables are not affected by the instrument for low percentiles of the $XBF/Asset$ distribution.

In a second group of tests we check the stability of our results when reconsidering the role of economically important "strategic regions" (SR) for our analysis (hereby we follow Carvalho, 2014). As discuss in Section 5.1, for SRs the exogeneity assumption behind the foreign funding shock could be arguably challenged if ΔXBF reflects a (pre-existent) drop in credit demand in regions that are important for banks' business. Moreover, SRs could be the target of fiscal transfers and liquidity support even within banks, what would lead us to estimate only a conservative "lower bound" of the true effect on credit and job creation. Alternatively, it could be argued that SRs are more vulnerable to the crisis provided that they have tighter direct (aggregate demand) ties with the U.S. economy, as long as they specialize in export products or commodities that lost value in the crisis.

We first address these concerns by replicating Equation (2) when excluding municipalities above the 90th percentile of the distribution in the following variables: the share of exports from municipality j to total country-level exports; the ratio of commodity exports (mining, agriculture, and food goods) to total country-level commodity exports; and the ratio of oil exports from municipality j to total country-level oil exports. These results, reported in Table A.13, do not alter our main findings.

Second, we re-run the model when dropping (i) the main economic centers (Sao Paulo, Rio de Janeiro, and Minas Gerais), and (ii) all federal states' capital municipalities. As shown in Table A.11, our results remain in place when these observations are excluded. Most importantly, the increase in the size of the relevant coefficients suggest that, if anything, SRs lead us to obtain conservative estimates of the pass-through of the shock on the real economy. This conclusion is also confirmed by re-estimating Equation (2) when adding an extended vector of municipality-level control variables. In this check we add to the log GDP the pre-crisis average monthly log change in WJC, the ratio of exports to GDP, the ratio of exports to the United States to GDP, the credit-to-GDP ratio, a dummy identifying small municipalities (those below the 25th percentile of pre-crisis log GDP), and the log of GDP per capita. These variables, which enter the model as pre-crisis averages, are expected to capture pre-existent aggregate demand trends and economic vulnerabilities that may confound the estimated effect of ΔXBF on macro aggregates. We do not find evidence of our results being affect by the inclusion of these control variables (see Table A.12).[12]

Despite these robustness tests, we may still be failing to capture the regional distribution of the country's export industry, which represents the sector that we would expect to be mostly directly linked to the global economy. This would be the case if the definition of strategic regions used above does not match the actual importance of exporting firms within regions. Therefore, we further test Equation (2) by excluding municipalities that we would expect to be mostly exposed to the global collapse in trade volumes and prices triggered by the crisis. We estimate Equation (2) for the subset of municipalities with a log change in exports (total and to the United States) around the crisis above the 10th percentile, as well for municipalities with a pre-crisis exports-to-GDP ratio below the 90th percentile of the respective distribution. These regressions confirm that our findings are not affected by specific trends in globally-exposed regions (see Table A.14).

[12]As a complementary check, we estimated a non-linear version of Equation (2) in which ΔXBF enters the model in an interaction term with variables measuring the economic importance of certain regions. First, we use dummies identifying municipalities in the main federal states and all federal states' capitals to test whether the effect depends on this cross-regional categorization. Second, we use as interaction terms the share of each municipality in the country's total exports, commodity exports (mining, agriculture, and food goods), and oil exports. These results, reported in Tables A.15 and A.16, confirm that our results are not spuriously driven by aggregate demand shocks.

In Table A.4 we report results showing that while the lending channel at the branch level mainly affects private-owned banks (in opposite to state-owned banks), foreign and U.S.-connected banks are especially sensible to the funding shock. To test whether these factors affect aggregate outcomes, we implement a test in which we separately estimate Equation (2) for municipalities with a large versus low market share of the aforementioned bank groups as defined by the sample median. This analysis, reported in Table A.18 using ΔWJC as the dependent variable, shows that the model renders statistically significant effects on job creation mostly in regions with a low presence of state-owned banks, with a large presence of foreign-owned banks (even though in this latter case the coefficient for ΔXBF is roughly the same across splits), or with a large market share of U.S.-connected banks.

We also implement a sample-split regression to shed light on whether the real effects can be linked to a general drop in credit supply or rather to a loss in commercial credit in particular. For this purpose, we split the municipalities' sample according to the median of the ratio of the log change in consumer credit to the log change in commercial credit around the crisis period. These growth rates are computed from aggregate credit in each segment at the municipal level. A higher value in this ratio indicates that consumer credit fell less (or grew more) in a given municipality than commercial credit. If corporate credit matters to explain firms' hiring dynamics, we would expect the effect on job creation to be stronger for higher values of the ratio. The results, reported in Table A.17, confirm that this is the case.[13]

Municipality-level results are also robust to rather mechanical adjustments to the empirical model (see Table A.19). For instance, the results hold when control variables enter the model as contemporaneous post-crisis averages, when standard errors are clustered at the municipality level, or when using fixed effects at a more broader macro-region level.[14] As an alternative specification, we use Equation (2) to estimate the log change in munic-

[13]In Table A.17 we also report a sample split test according to municipalities' GDP per capita. While our empirical approach does not suggest a specific conjecture on how this variable may affect our findings, the analysis is interesting as we could expect the social consequences of the real-economic adjustment triggered by the lending channel to differ over this dimension in the cross-section of municipalities. For instance, one could argue that in rather poor regions which likely display a low institutional quality, the adjustment in job creation may lead to an increase in informal labor. Also, poor institutional conditions may facilitate replacing formal work contracts by informal labor. Even though addressing this question would require regionally dis-aggregated data on informal labor that is not available in Brazil, looking at GDP per capita is interesting as a general test on whether the identified job creation effect is evenly distributed across poor versus rich regions or not. This sample split regressions show that the effect on ΔWJC is statistically significant in the subset of rather poor municipalities (even though the coefficients of interest are fairly similar in size). This result implies that we cannot rule out the hypothesis that the adjustment in job creation at least partially reflects a shift towards (cheaper) informal labor in affected municipalities.

[14]Macro-regions in Brazil correspond to 5 geographic regions defined for statistical purposes according to their economic, geographic, and cultural characteristics by the Brazilian Institute of Geography and Statistics (IBGE).

ipal GDP (between 2007 and 2009, end of year) and the log change in average net job creation by 1,000 inhabitants (i.e., hires minus layoffs). The pass-through of the lending channel to the real economy is further confirmed by these tests.

6 Conclusion

This paper documents how the characteristics of bank branches' funding networks affect the transmission of a capital flows reversal to the real economy. Using balance-sheet data connecting bank headquarters with their municipal branches in Brazil, we find robust evidence of a bank-branch lending channel of foreign funding shocks during the period of global capital flows reversals following the collapse of Lehman Brothers in September 2008. The lending channel increases in branches' funding fragmentation, measured as the distance to intra-bank funding networks and as how liquid those networks are.

Our analysis provides further evidence of the limited substitution of the lending channel across banks and firms. Municipality-level estimates show that both aggregate credit and job creation fell in municipalities hosting branches more exposed to the shock. While funding fragmentation matters also for aggregate credit and therefore for the substitution of funding within regions, job creation is only affected by fragmentation indirectly via the credit supply channel. We find, however, that the drop in job creation was partially offset when neighbouring regions are themselves less affected by the shock.

The effect of funding fragmentation may be related to other financial frictions previously identified in the literature on emerging countries finance, such as firms' limited pledgeability of cash flows (Khajam et al., 2010). Our analysis lends support to the idea that these microeconomic frictions may well interact with more general market structures in explaining the transmission of liquidity shocks originated abroad. Our findings suggest that the characteristics of bank branches' funding networks generate distributional effects across regions. This conclusion should be considered for the design and implementation of fiscal and monetary policies in times of financial distress, especially when the banking system is expected to intermediate those policy actions. Moreover, our findings provide support for policies promoting financial inclusion and the use of technology for banking transactions. These policies can reduce transaction costs and alleviate fragmentation frictions that impair capital reallocation.

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Figures and tables

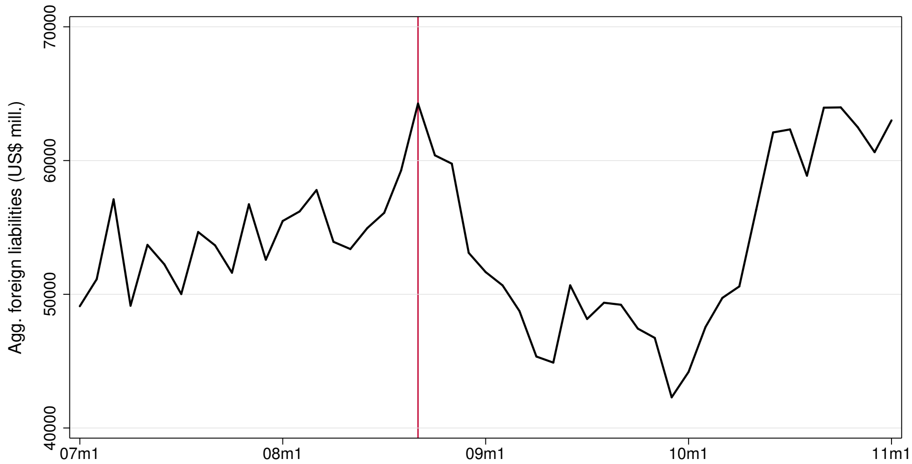

Figure 1 Time series of foreign funding

Notes: This figure displays the aggregated volumes of foreign funding, as measured by banks' total foreign interbank liabilities. The vertical line is set at September 2008, the month when the collapse of Lehman Brothers triggered a freeze in global interbank markets. Foreign funding is aggregated from the bank-level data in the baseline sample.

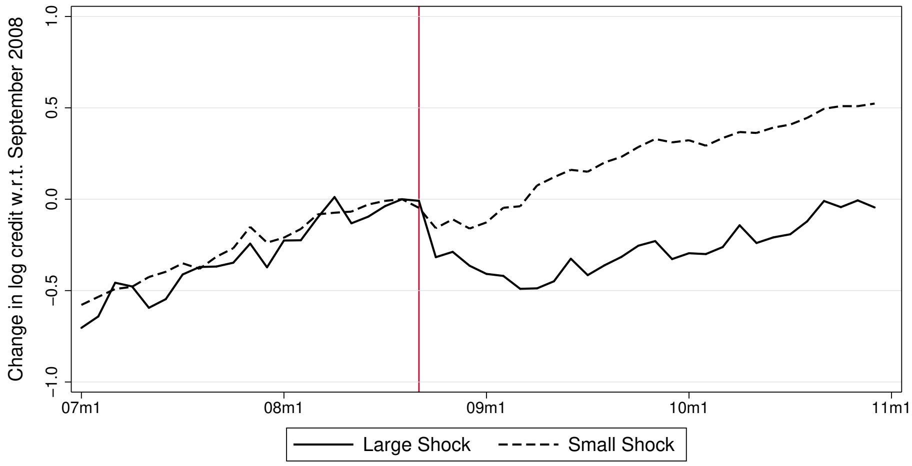

Figure 2 Bank lending channel

Notes: This figure illustrates the different pattern of credit growth followed by banks affected or not by a foreign funding shock after September 2008. The vertical line is set at September 2008, the month when the collapse of Lehman Brothers triggered a freeze in global interbank markets. The volume of outstanding credit is aggregated from the branch level data by bank and plotted as log first differences with respect to September 2008. Banks affected by a relatively large shock are those with a change in log foreign funding below the sample median.

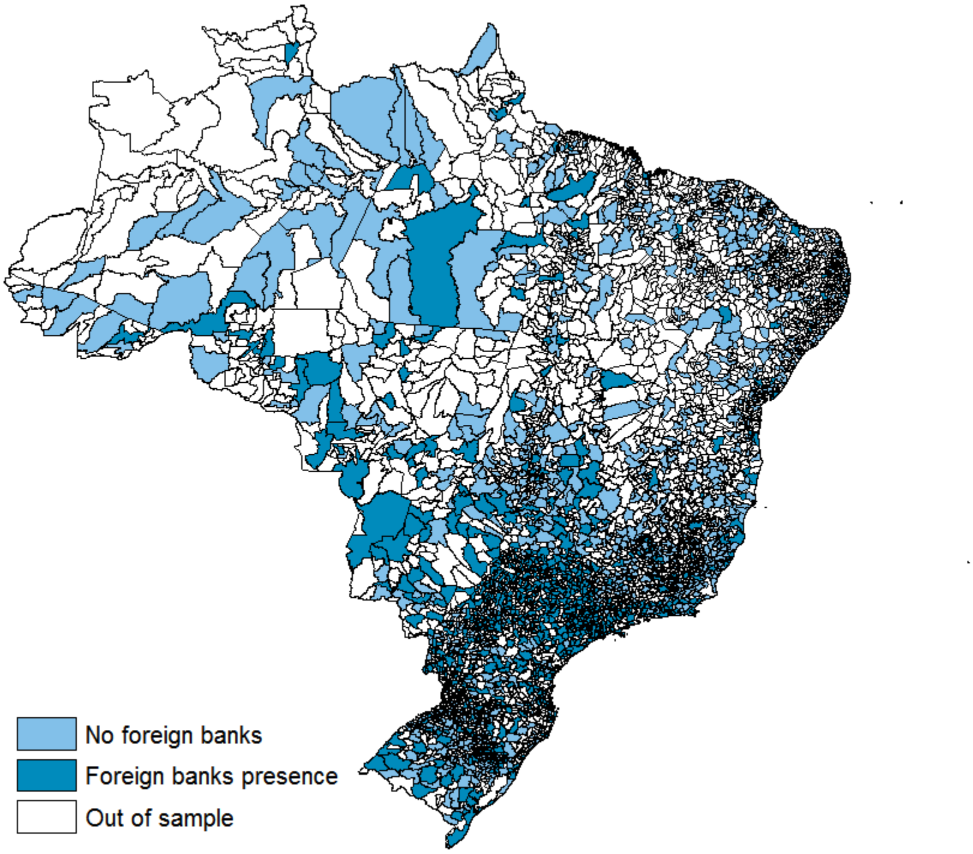

Figure 3 Geographical distribution of foreign banks

NOTES: This figure depicts the geographical distribution of foreign banks in the working sample. Regions in dark blue represent municipalities reporting some foreign bank activity through local bank branches between 2007 and 2010 and that are included in the sample. The regions in light blue are those in which no banking activity by foreign banks is reported within the working sample. Municipalities in white are the ones not included in the working sample. Foreign banks report activity in 714 out of 1,768 municipalities in the sample.

Table 1 DESCRIPTIVE STATISTICS – BRANCH SAMPLE

	Mean	SD	Min	Max	Shock-affected: large	low	Diff.
	I	II	III	IV	V	VI	VII
Var. of interest:							
ΔCredit	0.09	0.25	-0.37	0.67	0.05	0.13	-0.07*
ΔXBF	0.18	0.59	-1.38	1.68	-0.24	0.58	-0.82*
$XBF/Asset$	0.10	0.11	0.00	0.58	0.13	0.07	0.06
Parent-level:							
Size (log Assets)	3.34	6.78	0.38	29.34	2.75	3.04	-0.30
Capital ratio	0.14	0.09	0.02	0.44	0.16	0.12	0.04*
Liq. ratio	0.22	0.13	0.02	0.54	0.18	0.26	-0.09
Deposit base	0.37	0.18	0.00	0.75	0.31	0.42	-0.11
Credit risk	0.14	0.11	0.00	0.45	0.13	0.14	-0.01
State-owned	0.15	0.36	0.00	1.00	0.15	0.14	0.01
Foreign	0.39	0.49	0.00	1.00	0.50	0.29	0.21*
Branch-level:							
Size (log Assets)	0.66	4.39	0.00	39.83	2.13	0.18	1.95
Liq. ratio	0.10	0.11	0.00	0.49	0.09	0.12	-0.03
Internal fund.	0.32	0.28	0.00	0.80	0.33	0.32	0.01
Deposit base	0.39	0.29	0.04	0.96	0.36	0.43	-0.07
RoA	0.02	0.02	-0.02	0.08	0.01	0.02	-0.01
Pre-trend:							
Pre-ΔCredit	0.07	0.31	-0.69	1.17	0.06	0.02	0.04
Pre-ΔDeposits	0.05	0.22	-0.51	0.66	0.06	0.00	0.06

NOTES: This table reports descriptive statistics. The branch- and headquarter-level summary statistics are computed as pre-crisis values. Columns V and VI report the pre-crisis average for each variable within the groups of shock-affected and not-affected banks, respectively. Shock-affected banks are those reporting a value of ΔXBF below the sample median. Column VII shows the difference in means between affected and non-affected banks. The variables measuring Pre-crisis trends (Pre-trend) represent the average month-on-month change in log total credit and deposits (Pre-ΔCredit and Pre-ΔDeposits) at the branch level in the pre-crisis periods. * indicates whether the difference is significant by normalized differences (Imbens and Wooldridge, 2009), i.e., a value of larger than $|0.25|$. Variables are defined in Table 7 and winsorized at the 1st and 99th percentiles.

Table 2 BENCHMARK RESULTS – BRANCH LENDING CHANNEL

	Benchmark Model	Commercial Lending	Consumer Lending	Mortgage Lending	Leasing
Dependent variable	ΔCredit	ΔC&I	ΔCons	ΔMort	ΔLeas
	I	**II**	**III**	**IV**	**V**
ΔXBF	0.625**	0.382**	0.587**	0.202	0.029
	(0.288)	(0.180)	(0.244)	(0.152)	(0.084)
Parent-level					
Size (log Assets)	0.003	-0.004**	0.004	-0.002	-0.001
	(0.004)	(0.002)	(0.003)	(0.002)	(0.001)
Capital ratio	-2.145	-3.495***	-1.549	-2.538***	-0.268
	(1.478)	(0.898)	(1.198)	(0.830)	(0.391)
Liquidity ratio	-1.595**	-1.347***	-1.133*	-1.454***	-0.033
	(0.662)	(0.421)	(0.581)	(0.376)	(0.171)
Deposit base	-0.340	-0.937***	-0.192	-0.369	-0.062
	(0.383)	(0.311)	(0.318)	(0.229)	(0.046)
Credit risk	-2.331***	-0.868	-2.045***	-0.974***	-0.106
	(0.809)	(0.663)	(0.713)	(0.337)	(0.184)
State-owned	0.405***	0.198**	0.387***	0.128***	0.018
	(0.090)	(0.080)	(0.078)	(0.034)	(0.019)
Foreign	0.072	-0.018	0.131**	-0.039	0.005
	(0.069)	(0.058)	(0.059)	(0.041)	(0.009)
Branch-level					
Size (log Assets)	0.011***	0.012***	0.013***	0.024***	0.016
	(0.003)	(0.003)	(0.003)	(0.003)	(0.010)
Liquidity ratio	-0.459	-0.415	-0.383	0.173	0.399*
	(0.423)	(0.311)	(0.390)	(0.129)	(0.236)
Internal fund.	-0.182	-0.236	-0.074	-0.033	0.019
	(0.192)	(0.197)	(0.158)	(0.077)	(0.033)
Deposit base	-0.181	-0.182	-0.088	-0.014	0.010
	(0.189)	(0.192)	(0.159)	(0.080)	(0.028)
RoA	1.241**	0.253	1.480***	-0.643**	-0.111
	(0.588)	(0.428)	(0.515)	(0.286)	(0.087)
Obs.	6,632	6,632	6,632	6,632	6,632
R-squared	0.188	0.174	0.187	0.572	0.063

NOTES: This table reports the results of estimating Equation (1) for different specifications. In all regressions, the dependent variable is a measure of the change in log average outstanding credit between the post- and pre-crisis periods for specific credit segments. The pre-crisis period is between January 2007 and August 2008; the post-crisis period is between September 2008 and December 2010. Column I reports the baseline specification with municipality fixed effects from Equation (1) using total outstanding credit to compute the dependent variable (ΔCredit). Columns II to V replicate the estimation for the segments of commercial lending (ΔC&I), consumer lending (ΔC&I), mortgage lending (ΔC&I), and leasing (ΔLeas), respectively. All regressions include municipality FE. Standard errors (in parentheses) are clustered at the bank and municipality levels. For a detailed definition of all variables, see Table 7. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively.

Table 3 BRANCH-LEVEL EFFECTS – FRAGMENTATION

	Benchmkt.	Distance	Size Rank	Bank deposit	Regional deposit
	I	**II**	**III**	**IV**	**V**
ΔXBF	0.625**	0.078	0.640***	1.262***	1.714**
	(0.288)	(0.330)	(0.212)	(0.277)	(0.619)
ΔXBF x Distance		0.113**			
		(0.049)			
ΔXBF x Size Rank			-0.155*		
			(0.078)		
ΔXBF x Bank deposit				-1.663***	
				(0.466)	
ΔXBF x Reg. deposit					-1.351*
					(0.765)
Distance		0.058**			
		(0.024)			
Size rank			-0.102***		
			(0.024)		
Bank deposit				-1.570**	
				(0.661)	
Reg. deposit					0.357**
					(0.158)
Obs.	6,632	6,632	6,632	6,632	6,632
R-squared	0.351	0.189	0.201	0.222	0.212

NOTES: This table reports the results of estimating Equation (1) by adding an interaction term between ΔXBF and different variables measuring branches' financial fragmentation. In all regressions, the dependent variable is a measure of the change in log average outstanding credit between the post- and pre-crisis periods for specific credit segments. The pre-crisis period is between January 2007 and August 2008; the post-crisis period is between September 2008 and December 2010. Column I replicates the benchmark specification with municipality fixed effects reported in Table 2, Column I. Columns II to V report regressions adding the respective interaction term. The interaction variables include the distance to the closest branch of the same bank (Distance, Column II); the log of the size rank of a branch within its bank in a given federal state (Size rank, Column III); the ratio of aggregated deposits to total assets by all other branches different than branch i within its bank (BHC deposit, Column IV); and the average ratio of deposits to total assets by all branches different than branch i within a micro-region (Regional deposit, Column V). Regressions include municipality FE. Standard errors (in parentheses) are clustered at the bank and municipality level. For a detailed definition of all variables, see Table 7. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively.

Table 4 MUNICIPALITY-LEVEL EFFECTS – BENCHMARK

Dependent variable	ΔMcred		ΔJC		ΔWJC	
	OLS	FE	OLS	FE	OLS	FE
	I	**II**	**III**	**IV**	**V**	**VI**
ΔXBF	0.508***	0.602**	0.282**	0.345**	0.223**	0.318**
	(0.144)	(0.237)	(0.132)	(0.174)	(0.102)	(0.146)
log GDP	0.046***	0.036***	0.011	-0.004	0.009*	-0.004
	(0.006)	(0.009)	(0.007)	(0.008)	(0.005)	(0.006)
Parent-level						
Size (log Assets)	0.040***	0.043***	-0.010*	-0.010*	-0.011**	-0.009**
	(0.007)	(0.009)	(0.006)	(0.006)	(0.005)	(0.005)
Capital ratio	3.894***	4.810***	0.371	0.114	-0.435	-0.424
	(0.923)	(1.207)	(0.993)	(1.127)	(0.767)	(0.867)
Liquidity ratio	-0.333	-0.591	-1.596***	-1.797**	-1.234***	-1.589***
	(0.575)	(1.031)	(0.515)	(0.707)	(0.412)	(0.599)
Deposit base	0.149	0.036	-0.495	-0.524	-0.447	-0.580*
	(0.469)	(0.502)	(0.425)	(0.412)	(0.341)	(0.337)
Credit risk	-3.095***	-3.474***	0.791	0.616	0.756*	0.476
	(0.576)	(0.705)	(0.560)	(0.579)	(0.430)	(0.449)
State-owned	0.757***	0.697***	-0.012	-0.121	-0.054	-0.124
	(0.104)	(0.128)	(0.088)	(0.103)	(0.067)	(0.080)
Foreign	0.764***	0.724***	-0.329***	-0.174*	-0.274***	-0.156*
	(0.112)	(0.141)	(0.093)	(0.100)	(0.073)	(0.081)
Branch-level						
Size (log Assets)	-1.228***	-1.052***	-0.372	0.151	-0.210	0.230
	(0.300)	(0.374)	(0.290)	(0.311)	(0.224)	(0.253)
Liquidity ratio	2.270***	1.926***	1.899***	-0.013	1.231***	0.004
	(0.369)	(0.452)	(0.439)	(0.418)	(0.287)	(0.305)
Internal fund.	-0.188	-0.235	0.006	0.043	0.015	0.076
	(0.139)	(0.163)	(0.128)	(0.111)	(0.090)	(0.083)
Deposit base	0.020	-0.037	-0.014	-0.024	-0.011	0.015
	(0.147)	(0.171)	(0.139)	(0.122)	(0.095)	(0.093)
RoA	-2.534***	-2.307***	-0.277	0.317	-0.249	0.231
	(0.593)	(0.764)	(0.493)	(0.494)	(0.365)	(0.370)
Obs.	1,768	1,640	1,768	1,640	1,768	1,640
R-squared	0.240	0.176	0.054	0.007	0.045	0.007

NOTES: This table reports the results of estimating Equation (2) for different real economic outcomes at the municipality level. This data set is aggregated at the municipality level from the branch-level sample. Standard errors are clustered at the micro-region level. The real outputs considered are the change in log aggregated outstanding credits (ΔMcred, Columns I and II), the change in the log number of new contracts (ΔJC, Columns III and IV), and the change in log job creation per 1,000 inhabitants (ΔWJC, Columns V and VI). For each dependent variable the table reports the results of simple OLS regressions (Columns I, III and V) and regressions with fixed effects at the micro-region level (Columns II, IV and VI). Standard errors are clustered at the micro-region level. For a detailed definition of all variables see Table 7. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively.

Table 5 Municipality-level credit growth – Fragmentation

	Dependent variable: ΔMcredit				
Municipality trait:	Distance	Size rank	BHC deposit	Reg. Deposit	ΔXBF-reg
	I	**II**	**III**	**IV**	**V**
ΔXBF	0.413*	0.771***	9.062***	2.194**	0.636***
	(0.243)	(0.265)	(3.428)	(0.855)	(0.197)
ΔXBF x Trait	0.008*	-0.003**	-9.765**	-1.776**	-1.515
	(0.005)	(0.001)	(4.122)	(0.870)	(1.354)
Trait	-0.001	0.001	7.046***	1.496***	0.037
	(0.002)	(0.001)	(1.630)	(0.417)	(0.495)
log GDP	0.033***	0.044***	0.026**	0.043***	0.041***
	(0.009)	(0.010)	(0.012)	(0.008)	(0.006)
Parent-level					
Size (log Assets)	0.047***	0.043***	0.042**	0.041***	0.040***
	(0.009)	(0.008)	(0.017)	(0.008)	(0.007)
Capital ratio	5.293***	4.623***	11.703***	5.940***	3.786***
	(1.188)	(1.126)	(3.895)	(1.442)	(0.940)
Liquidity ratio	-0.496	-0.674	-5.146***	-0.394	-0.291
	(1.008)	(1.035)	(1.717)	(0.950)	(0.576)
Deposit base	0.301	0.022	-1.507*	-0.259	0.265
	(0.505)	(0.500)	(0.790)	(0.507)	(0.465)
Credit risk	-3.655***	-3.375***	-2.434***	-2.558***	-3.068***
	(0.710)	(0.699)	(0.860)	(0.675)	(0.579)
State-owned	0.703***	0.663***	0.461***	0.725***	0.742***
	(0.127)	(0.129)	(0.124)	(0.116)	(0.104)
Foreign	0.721***	0.694***	0.862***	0.752***	0.750***
	(0.140)	(0.143)	(0.136)	(0.128)	(0.112)
Branch-level					
Size (log Assets)	-1.136***	-1.190***	-0.401	-1.197***	-1.150***
	(0.365)	(0.382)	(0.512)	(0.363)	(0.300)
Liquidity ratio	1.741***	1.786***	3.136***	1.669***	2.108***
	(0.457)	(0.508)	(0.648)	(0.468)	(0.359)
Internal fund.	-0.246	-0.215	-0.207	-0.205	-0.201
	(0.171)	(0.165)	(0.192)	(0.171)	(0.140)
Deposit base	-0.051	-0.029	-0.066	-0.073	0.005
	(0.177)	(0.173)	(0.201)	(0.175)	(0.147)
RoA	-2.338***	-2.395***	-2.151***	-2.460***	-2.433***
	(0.746)	(0.794)	(0.777)	(0.766)	(0.571)
Obs.	1,640	1,640	1,640	1,640	1,640
R-squared	0.184	0.176	0.167	0.198	0.246

Notes: This table reports the results of estimating Equation (2) with an interaction term between ΔXBF and municipality-level variables measuring the characteristics of branches' funding network structure. In all regressions, the dependent variable is a measure of the change in log average outstanding credit (aggregated at the municipal level) between the post- and pre-crisis periods (ΔMCred). The variable ΔXBF is interacted with the following variables (i.e., Trait): the log distance to the closest branch (Column I); the log of the size rank within the respective federal state (Column II); the ratio of deposits to assets at the bank level excluding branch i, j (Column III), the average ratio of deposits to assets by all other branches of the same bank i within the micro-region in which municipality j is located (Column IV); and the average of ΔXBF within the micro-region in which municipality j is located (Column V). The variables Trait in Columns I to IV and the control variables are measured as market-share-weighted averages of all branches within municipality j. Regressions include micro-region FE. Standard errors are clustered at the micro-region level. For a detailed definition of all variables, see Table 7. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively.

Table 6 MUNICIPALITY-LEVEL JOB CREATION GROWTH – FRAGMENTATION

	Dependent variable: ΔWJC				
Municipality trait:	Distance	Size rank	BHC deposit	Reg. Deposit	ΔXBF-reg
	I	**II**	**III**	**IV**	**V**
ΔXBF	0.305***	0.252**	0.317*	0.321*	-0.176
	(0.109)	(0.114)	(0.192)	(0.183)	(0.233)
ΔXBF x Trait	-0.005	0.000	-0.212	-0.155	4.295**
	(0.003)	(0.001)	(0.257)	(0.201)	(1.898)
Trait	0.001	-0.001*	-0.018	-0.020	-1.671**
	(0.001)	(0.000)	(0.084)	(0.067)	(0.701)
log GDP	0.009	0.004	0.007	0.007	0.009
	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
Parent-level					
Size (log Assets)	-0.011**	-0.012***	-0.010**	-0.010**	-0.013***
	(0.005)	(0.004)	(0.005)	(0.005)	(0.004)
Capital ratio	-0.625	-0.824	-0.532	-0.486	-0.820
	(0.773)	(0.770)	(0.782)	(0.799)	(0.726)
Liquidity ratio	-1.241***	-1.392***	-1.052**	-1.171***	-1.225***
	(0.414)	(0.401)	(0.416)	(0.413)	(0.415)
Deposit base	-0.508	-0.640*	-0.244	-0.342	-0.549*
	(0.343)	(0.345)	(0.350)	(0.348)	(0.333)
Credit risk	0.828*	0.914**	0.758*	0.643	0.874**
	(0.436)	(0.421)	(0.428)	(0.433)	(0.426)
State-owned	-0.055	-0.133*	-0.073	-0.073	-0.073
	(0.067)	(0.068)	(0.066)	(0.067)	(0.066)
Foreign	-0.267***	-0.244***	-0.272***	-0.272***	-0.272***
	(0.074)	(0.073)	(0.073)	(0.073)	(0.075)
Branch-level					
Size (log Assets)	-0.237	-0.029	-0.094	-0.011	-0.272
	(0.226)	(0.230)	(0.226)	(0.226)	(0.223)
Liquidity ratio	1.258***	0.778***	0.990***	1.031***	1.234***
	(0.283)	(0.271)	(0.285)	(0.286)	(0.285)
Internal fund.	0.010	0.023	0.025	0.025	0.029
	(0.091)	(0.092)	(0.090)	(0.088)	(0.091)
Deposit base	-0.019	-0.002	0.000	0.032	-0.006
	(0.096)	(0.098)	(0.096)	(0.095)	(0.096)
RoA	-0.212	-0.234	-0.170	-0.185	-0.348
	(0.363)	(0.357)	(0.360)	(0.361)	(0.364)
Obs.	1,640	1,640	1,640	1,640	1,640
R-squared	0.046	0.064	0.054	0.052	0.053

NOTES: This table reports the results of estimating Equation (2) with an interaction term between ΔXBF and municipality-level variables measuring the characteristics of branches' funding network structure. In all regressions, the dependent variable is a measure of the change in log job creation by 1,000 inhabitants (aggregated at the municipal level) between the post- and pre-crisis periods (ΔWJC). The variable ΔXBF is interacted with the following variables (i.e., Trait): the log distance to the closest branch (Column I); the log of the size rank within the respective federal state (Column II); the ratio of deposits to assets at the bank level excluding branch i, j (Column III), the average ratio of deposits to assets by all other branches of the same bank i within the micro-region in which municipality j is located (Column IV); and the average of ΔXBF within the micro-region in which municipality j is located (Column V). The variables Trait in Columns I to IV and the control variables are measured as market-share-weighted averages of all branches within municipality j. Regressions include micro-region FE. Standard errors are clustered at the micro-region level. For a detailed definition of all variables, see Table 7. *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively.

Table 7 VARIABLES DEFINITION

Variable	Definition	Unit
Bank vars:		
ΔXBF	Change in log average foreign liabilities in the post- minus pre-crisis periods.	Rate
$XBF/Asset$	Ratio of interbank funding originated outside Brazil and total assets.	Rate
Size (log Assets)	Total size of a bank's balance sheet computed as the log of total assets in millions of USD.	log
Capital ratio	Ratio of equity to total assets.	Rate
Liquidity ratio	Ratio of liquid to total assets at the parent-bank level. Liquid assets are defined as cash holdings.	Rate
Deposit base	Ratio of sight plus saving deposits to total assets.	Rate
Credit risk	Ratio of non-performing to total outstanding credit. Non-performing credits are loans reporting some delay in its re-payment record. The ratio is computed by dividing the volume of non-AA-rated loans to total credit.	Rate
Foreign	Dummy variable equal to 1 if a bank is owned by a foreign financial institution and 0 otherwise.	0/1
State-owned	Dummy variable equal to 1 if a bank is owned by the Brazilian state and 0 otherwise.	0/1
N-con	Dummy variable equal to one for banks without a related institution (headquarter, affiliate, or a bank of the same banking group) located in the United States as of 2007.	0/1
$Finv$	Dummy variable equal to one for banks with a ratio of foreign investments to assets above the median of the pre-crisis average.	0/1
Branch vars:		
ΔCredit	Change in log average outstanding credit in the post- minus pre-crisis periods.	Rate
ΔC&I	Change in log average outstanding commercial credit in the post- minus pre-crisis periods.	Rate
ΔCons	Change in log average outstanding consumer credit in the post- minus pre-crisis periods.	Rate
ΔMort	Change in log average outstanding mortgage credit in the post- minus pre-crisis periods.	Rate

NOTES: This table provides a description of the main variables used for the empirical analysis reported in the paper. The variables are divided into the following categories: parent bank variables (Bank vars.), municipal branch variables (Branch vars.), fragmentation variables (Frag. vars.), and municipality variables (Municipality vars.).

Table 7 Variables definition (continued)

Variable	Definition	Unit
Branch vars:		
Pre-ΔCredit	Average month-on-month change in log outstanding credit in the pre-crisis period.	Rate
Pre-ΔDeposits	Average month-on-month change in log deposits in the pre-crisis period.	Rate
Size (log Assets)	Total size of a branch's balance sheet computed as the log of total assets measured in millions of USD.	log
Liquidity ratio	Ratio of liquid to total assets at the branch-bank level. Liquid assets are defined as cash holdings.	Rate
Internal fund.	Ratio of internal liabilities to total assets at the branch-bank level. Internal liabilities are defined as liabilities vis-à-vis correspondent entities within the same bank.	Rate
Deposits base	Ratio of sight plus saving deposits to total assets.	Rate
RoA /	Ratio of net income to assets.	Rate
Frag. vars:		
Distance	Geographical distance to the closest branch of the same bank in log km. Distances are measured using municipalities' geographical centroids.	log
Size rank	Log of size rank of a branch within its bank in a given federal state. Rank 1 represents the smallest branch.	log
BHC deposit	Ratio of aggregated deposits (sight plus saving) to total assets by all other branches of a given bank (i.e., bank holding company, BHC), excluding branch i.	Rate
Reg. deposit	Average deposit base by all branches different than branch i within a micro-region.	Rate
Municipality vars:		
ΔMcred	Change in log average aggregated outstanding credit by all branches within a municipality in the post- minus pre-crisis periods. The aggregation is from the underlying monthly balances.	Rate
ΔJC	Change in average log job creation as measured by new job contracts in the post- minus pre-crisis periods.	Rate
ΔWJC	Change in average log job creation as measured by new job contracts by 1,000 inhabitants in the post- minus pre-crisis periods.	Rate

Table 7 VARIABLES DEFINITION (CONTINUED)

Variable	Definition	Unit
Municipality vars:		
ΔWNJC	Change in average log net job creation as measured by net new job contracts by 1,000 inhabitants in the post- minus pre-crisis periods. Net job creation is computed by subtracting the terminated contracts from the number of new contracts reported per month in a given municipality.	Rate
ΔGDP	Log change in municipal GDP between 2009 and 2007 (end of year data)	Rate
Main regions	Dummy variable equal to one for municipalities within the Brazilian federal states (UF) of Sao Paolo, Rio de Janeiro, and Minas Gerais.	0/1
Population (Pop.)	Log of municipal population as of 2007.	log
AG size	Pre-crisis average of the log of aggregate total assets at the municipal level. Total assets are added-up per month considering all branches operating in a given municipality.	log
No branches	Pre-crisis average of the log the number of banks operating in a given municipality.	log
Av mkt share	Pre-crisis average of the (monthly) market share of all branches operating in a municipality.	Rate
HHI	Pre-crisis average of the Herfindahl Index computed as the municipal level. The index is computed from branches' market shares using as a reference their total assets. A large index proxies for a more concentrated market.	Rate
Log GDP	Log of the pre-crisis (2007) municipal GDP.	log
Log GDP pc	Pre-crisis (2007) log of the municipal GDP per capita.	log
UF cap.	Dummy equal to 1 for municipalities defined as the capital city of their respectivee federal states (Unidades Federativas, UF).	0/1

Table 7 Variables definition (continued)

Variable	Definition	Unit
Municipality vars:		
av. ΔWJC	Pre-crisis average of the monthly change in log job creation by 1,000 inhabitants at the municipal level.	Rate
log Credit/GDP	Ratio of the pre-crisis average aggregate credit to GDP at the municipal level	log
$Exp_j/T.Exp$	Average share of pre-crisis total exports from municipality j in Brazilian total exports. The average is computed from monthly pre-crisis series.	Rate
$Comm_j/T.Comm$	Average share of pre-crisis commodity exports from municipality j in Brazilian total commodity exports. Commodities are defined as export products in the mining, oil, and agricultural industries. The average is computed from monthly pre-crisis series.	Rate
$Oil_j/T.Oil$	Average share of pre-crisis oil-industry exports from municipality j in Brazilian total oil-industry exports. The average is computed from monthly pre-crisis series.	Rate
Export/GDP	Average pre-crisis ratio of total exports to GDP per municipality	Rate
Δ^-Exports	Log change in average total exports between the pre- and post-crisis periods. The average is computed from monthly pre-crisis series.	Rate
Δ^-Exports US	Log change in average total exports to the United States between the pre- and post-crisis periods. The average is computed from monthly pre-crisis series.	Rate

Halle Institute for Economic Research –
Member of the Leibniz Association

Kleine Maerkerstrasse 8
D-06108 Halle (Saale), Germany

Postal Adress: P.O. Box 11 03 61
D-06017 Halle (Saale), Germany

Tel +49 345 7753 60
Fax +49 345 7753 820

www.iwh-halle.de

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