Essays in Empirical Macro-Finance

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Abstract

Do financial intermediaries influence firms' (real and financial) outcomes? Are there interactions with monetary and macroprudential policy? This thesis tackles those questions from multiple perspectives. The first chapter asks whether monetary policy shocks alter the maturity structure of US corporate debt. I find that looser monetary conditions lengthen non-financial firms' debt maturity, especially among very big corporations and driven by increased risk-taking by bond-market investors. The second and third chapter analyze the effects of prudential capital controls on corporate debt and real outcomes, looking at the recent experience of Colombia. Capital controls are shown to reduce corporate debt during a boom - either directly, or indirectly, i.e. by amplifying the contractionary effects of higher policy rates on banks' credit supply - thereby improving firm performance during the bust.

Keywords: Monetary Policy, Corporate Debt, Maturity, Reach-for-yield, Financial Frictions, Capital Controls, Macroprudential Policy, Carry trade, Credit Supply, Risk-taking, FX-debt, Real Effects, Capital Inflows.

Resumen

Influyen los intermediarios financieros en los resultados (reales y financieros) de las empresas? ¿Existen interacciones entre la política monetaria y macroprudencial? Esta tesis aborda estas cuestiones desde diferentes ángulos. El primer capítulo estudia si los *shocks* de política monetaria alteran la estructura de vencimiento de la deuda corporativa estadounidense. Los resultados apuntan a que unas condiciones monetarias más flexibles alargan el vencimiento de la deuda de las empresas no financieras, especialmente entre las corporaciones muy grandes y están impulsadas por un incremento del riesgo asumido por los inversores del mercado de bonos. El segundo y tercer capítulo analizan los efectos de los controles prudenciales de capital sobre la deuda corporativa y los efectos reales, basándose en la experiencia reciente de Colombia. Se demuestra que los controles de capital reducen la deuda corporativa durante un auge, mejorando así el desempeño de las empresas durante la crisis, ya sea directamente o indirectamente, es decir, amplificando los efectos contractivos de un incremento de los tipos de interés oficiales sobre la oferta crediticia de los bancos.

Palabras clave: Política Monetaria, Deuda Corporativa, Vencimiento, Fricciones Financieras, Controles de Capital, Política Macroprudencial, Carry Trade, Oferta Crediticia, Asunción de riesgos, Deuda en Moneda Extranjera, Efectos Reales, Flujos de Capital.

Preface

This thesis analyzes whether financial intermediation, by both banks and other financial investors, influences non-financial companies (real and financial) outcomes. The question is tackled from multiple perspectives and the results validate the hypothesis that financing frictions at the level of the firm - and of its lenders as well - can have a significant impact on the transmission of financial shocks to the real economy. Moreover, analyzing different macroeconomic policies, including monetary policy and macroprudential shocks, the different chapters of this thesis also reveal that financial intermediaries and bond-market investors alike have an important role in the the pass-through of such policy interventions to non-financial companies (NFCs hereafter).

In detail, the first chapter asks whether monetary policy shocks alter the maturity structure of US corporate debt. This question is important in that debt maturity is key for explaining NFCs' ability to withstand negative shocks, as documented by a large empirical literature on the last Great Financial Crisis of 2008-2009 and on the current COVID-19 pandemic. The analysis exploits conventional firm-level and macroeconomic datasets for the US economy, such as Compustat, Mergent FISD and CRSP Mututal Funds data. The results indicate that an expansionary monetary policy (interest rate) shock overall lengthens debt maturity for the US non-financial corporate sector. Furthermore, this effect entirely concentrates among very large firms, i.e. those in the top asset-size quartile of the respective industry-level distribution. These findings are somewhat surprising: in fact, interest rate shocks are typically expected to alter the capital structure of smaller NFCs. However, a simple model featuring firm-level financing frictions and reach for yield by financial investors rationalizes the empirical facts. When the interest rate goes down, financial investors increase risk-taking, thereby expanding their demand for long-term debt securities. Larger firms with easier access to bond financing accommodate the upward demand shift and obtain a substantial discount in their financing costs. Further empirical evidence on the response of corporate bonds' issuance by large companies and holdings by mutual funds validates such mechanism.

The second and third chapter analyze the effects of prudential capital controls on corporate debt and real outcomes. This policy is at the center stage of an important debate (in both policy and academic circles) about the potential benefits of temporary restrictions to capital mobility, which might improve financial stability and/or strenghten the effectiveness of national macroeconomic policies. Both papers in this thesis exploit the recent pre-Crisis experience of Colombia, which introduced a tax on foreign debt in 2007, and take advatange of rich administrative and confidential loan-level and firm-level datasets. A first article (i.e., the second chapter of the thesis) shows that capital controls reduce credit to NFCs during a boom, in particular for NFCs with weaker relationships with domestic banks, which cannot substitute the forgone credit from abroad with financing from their national lenders. For these firms, exposure to capital controls is associated to a cut in imports during the boom. However, there is also evidence of prudential benefits, namely an increase in exports during the subsequent bust, larger for riskier and/or financially constrained firms, which otherwise tend to be more negatively impacted by high levels of leverage during major financial downturns.

The third chapter of the thesis exploits the same policy and data to understand whether capital controls and/or other domestic macroprudential measures (such as an increase in reserve requirements on domestic deposits) strenghten the transmission of monetary policy rates to credit supply. The findings suggest that, under capital mobility, the pass-through is weakened by a carry-trade lending strategy by banks. As a matter of fact, in reaction to an increase in the local policy rate, the interest rate differential against the financial center (i.e., the US) goes up. Banks respond by increasing their borrowing in cheap foreign currency and expanding their credit supply in domestic currency, thereby gaining (at least) the interest rate differential. Capital controls tax foreign debt and halt the carry, contributing to re-establishing a more negative relation between the monetary policy rate and credit supply. Differently, domestic macroprudential measures cut credit supply directly, rather than through their influence on the transmission of monetary policy rates. Additionally, we find that reliance on foreing funding and on domestic deposits are strongly negatively correlated across banks' balance sheets, so that banks more impacted by capital controls are less exposed to domestic macroprudential measures, and vice versa. Overall, this study establishes a prudential Tinbergen rule: taming a boom driven by both foreign and domestic liquidity requires two policy instruments, capital controls and domestic macroprudential measures.

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1. MONETARY POLICY AND CORPORATE DEBT MATURITY

Joint with Luigi Falasconi and Janko Heineken

1.1 Introduction

High corporate indebtedness is a major source of vulnerability for many Advanced Economies (IMF 2019*a*,*b*, Kaplan 2019).[1] In this respect, the maturity structure is a key feature of corporate debt, influencing firms' reaction to both real and financial shocks (Almeida et al. 2009, Duchin et al. 2010, He & Xiong 2012*b*, Kalemli-Ozcan et al. 2018, Jungherr & Schott 2020*b*, Chen et al. 2020). The ongoing Covid-19 crisis is not an exception: in fact, a higher share of maturing obligations has been found to depress non-financial firms' stock returns (Fahlenbrach et al. 2020) and to limit their access to capital markets (Halling et al. 2020) during the most acute phases of the pandemic. Understanding whether and how monetary policy affects the maturity structure of corporate debt is therefore of utmost importance, as it allows to gauge a potentially relevant bearing of central banks' policy on firms' risk. However, up to our knowledge, existing studies do not systematically explore this question.

We fill the gap by investigating the influence of the interest rate policy by the FED on the maturity structure of the US corporate debt sector. Our focus is over the period 1990-2017 and our empirical exercise exploits: i) different measures of endogenous and exogenous - variation of the Effective Fed Funds Rate (EFFR); ii) a large variety of time-series, firm-level and security-level datasets.

We find that, at the aggregate level, a reduction of the EFFR *lengthens* corporate debt maturity (i.e., increases the share of total debt with maturity above 1-year). The effect is both statistically and economically significant. A 25 basis points (b.p.) descent in the EFFR triggers a persistent jump in the share of long-term (LT) debt, amounting to roughly 0.42 percentage points (p.p) one year after the shock. For comparison, the average quarterly growth rate of the share of LT debt equals 0.15 p.p.. Next, we look at quarterly balance sheets of US listed firms and find that very large companies - namely those in the top-quartile of their respective industry-wide asset-size

[1] See Giroud & Mueller (2017, 2018) on how high firms' leverage boosts business-cycle fluctuations.

distribution - are responsible for the observed aggregate patterns, whereas smaller firms do not adjust.

We explain such findings through a parsimonious model combining financial frictions due to moral hazard (Holmström & Tirole 1998, 2000) and short-termist, yield-oriented investors (Hanson & Stein 2015), who care about current portfolio yield on top of expected returns and rebalance their portfolios toward LT debt securities when the policy rate descents. The demand shift decreases bond yields, but only large and unconstrained companies can take advantage by issuing LT bonds.

Our model delivers predictions aligned to our aggregate and cross-sectional evidence, and we empirically test its mechanism. We find that yield-oriented corporate bonds mutual funds increase their holdings of corporate bonds (as compared to other funds) when the policy rate goes down, while also tilting their holdings towards longer-term debt securities. Moreover, large companies' likelihood of issuing LT bonds jumps more strongly and the coupon rate reacts with a stronger decline. This suggests that relative fluctuations in bonds issuance by large companies are demand-driven.

Our main contribution is to provide systematic evidence on the relation between monetary policy and the maturity structure of the debt of US non-financial corporations (NFCs). Other papers condition the impact of monetary policy shocks on the ex-ante heterogeneity in firm debt maturity (Ippolito et al. 2018, Jungherr & Schott 2020*a*). Differently, we document that the maturity structure of debt endogenously responds to monetary policy shocks. Furthermore, our paper adds to a novel set of studies documenting the role of yield-oriented investors in the transmission of monetary policy (see, e.g., Hanson & Stein 2015, Di Maggio & Kacperczyk 2017, Daniel et al. forthcoming, Lian et al. 2019). We innovate by linking the relation between monetary policy shocks and bond issuance to yield-oriented investors and showing the implications for firms' debt maturity structure.

The rest of this introduction is divided into two parts. First, we provide a detailed preview of the paper. Second, we discuss more thoroughly the related literature and contrast it with our paper.

DETAILED PREVIEW OF THE PAPER

We investigate two main research questions. First, we ask whether monetary policy has any impact on the maturity structure of corporate debt. Second, we verify eventual cross-sectional differences across companies in such relation. Our focus rests on conventional interest rate policy by the FED over the period 1990-2017 and we limit our attention to the US corporate sector.

We use three different variables for capturing the FED interest rate policy. The baseline exercises use the raw quarterly variation of the EFFR, an endogenous measure of changes in the monetary policy stance displaying large persistence over tightening and loosening cycles (Adrian et al. 2010). Still, this measure has a direct "real-world" impact on firms' financing cost, and hence we test its influence on the debt maturity structure while controlling for other correlated macroeconomic variables such as GDP growth and inflation rate. Importantly, we also test the robustness of our results to two alternative exogenous measures of (high-frequency) interest rate shocks, borrowed from Gürkaynak et al. (2005) and Jarocinski & Karadi (2020). Our remaining data come from various sources. To start with, the time-series analysis of corporate debt maturity is based on quarterly data from FED Flows of Funds. Following, among others, Greenwood et al. (2010), we build the share of LT debt as the ratio between corporate debt with maturity above 1-year and total corporate debt. At the firm-level, we apply an identical measure, retrieved from Compustat quarterly financial data of US listed companies. To investigate our model-based mechanism, we get data on the universe of LT-bond issuance from Mergent FISD and access information on corporate bond mutual funds' holdings from the CRSP Survivor-Bias-Free US Mutual Fund dataset.

We study the aggregate evolution of LT debt by looking at the change in the LT debt share from the Flows of Funds, which mean equals 0.15 p.p. on a quarterly basis. We employ local projections (Jordà 2005) in a model augmented with other lagged macroeconomic controls. We find that a reduction of the EFFR lengthens corporate debt maturity, i.e., it expands the share of debt with maturity above 1-year. The effect is both statistically and economically significant. A 25 basis points (b.p.) descent in the EFFR triggers a persistent jump in the share of LT debt, amounting to roughly 0.42 p.p. one year after the shock. The effect peaks up three years after the shock. Results are robust to using exogenous monetary policy shocks.[2]

Next, we test whether such effect is heterogeneously distributed across companies. For this purpose, we use quarterly balance sheet data for US listed firms from Compustat. The key layer of heterogeneity is firm-size (approximated through total assets), inversely related to the intensity of financial constraints. In particular, we sort firms' in quartiles of the (lagged) industry-level (3-

[2] The effect has similar magnitude to that described for the EFFR variations, though it is less persistent - a difference we impute to the significant autocorrelation characterizing the raw EFFR series.

digit SIC classification) size distribution. We again apply local projections to analyze the dynamic response of the share of long-term debt, in a setting akin to Ottonello & Winberry (forthcoming) and Jeenas (2018). Our interest falls on the interaction between the change of the EFFR and a dummy identifying large companies, i.e. those in the top-quartile of their industry level size-distribution. This allows us to saturate the model with firm and industry*year-quarter fixed effects, controlling for firm-level time-invariant heterogeneity and industry-wide time-varying shocks, respectively. Also, we horserace this channel against other relevant firm balance-sheet items (fully interacted with the variation of the policy rate). Our results show that large companies adjust more, that is, they increase (decrease) their share of LT debt relatively more in response to a descent (jump) in the policy rate. Moreover, running separate regressions for companies in different quartiles of the asset-size distribution, we find that smaller companies' debt maturity is generally not responsive to monetary policy. Using exogenous interest rate shocks produces comparable results.

Interestingly, it is not easy to rationalize our results according to standard models of monetary policy transmission to firms, which, in general, do not consider the debt maturity structure.[3] On the other hand, the corporate finance literature does not directly focus on the policy rate, but rather on the term-spread, delivering counterfactual predictions relative to our findings.4 Hence, we propose a theory that can account for our aggregate and cross-sectional empirical facts.

We augment a standard model with short and long-term debt and financing frictions due to moral-hazard (Holmström & Tirole 1998, 2000) with the presence of short-termist, yield-seeking investors (Hanson & Stein 2015). Such investors are assumed to take long-short positions and care about current portfolio yield and not just expected returns. This modeling assumption reflects short-termist incentives of important classes of investors which, for instance, report each quarter to the stock market and therefore care about current yields on top of total expected returns. As a

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[3] A classical literature on the credit channel (Bernanke & Gertler 1995) of monetary policy exploits either frictions at the level of the company (e.g. Gertler & Bernanke 1989, Bernanke et al. 1999, Iacoviello 2005, Christensen & Dib 2008, Christiano et al. 2014) or at the level of the firms' lenders, typically banks (e.g. Kashyap & Stein 1994, Adrian & Shin 2010, Gertler & Karadi 2011, Borio & Zhu 2012, Dell'Ariccia et al. 2014). Under both paradigms, an interest rate cut relaxes credit standards, typically proxied by loan volume and rate, with greater relative benefits for small and constrained companies. Heuristically, such models would likewise predict a greater expansion of debt maturity for smaller companies, and hence may contradict our cross-sectional evidence. More novel contributions - including e.g. Ottonello & Winberry (forthcoming), Ozdagli (2018) - highlight how larger and less constrained NFCs may respond more to monetary policy shocks but neglect debt maturity.

[4] The reasoning goes as follows: a policy rate cut widens the term-spread (a stable relation documented by, e.g., Adrian & Shin 2010 and which we also verify in our sample) and hence increases the relative convenience of short-term debt issuance.

result, in reaction to a policy rate cut, they rebalance their portfolios toward longer term debt in an effort to keep their portfolio yield up, ultimately creating buying pressure on the price for LT debt. The boost in demand for LT debt is accommodated by larger NFCs, for which borrowing constraints are not binding.[5]

The model delivers predictions in line with both time-series and cross-sectional evidence and dovetails nicely with both large corporations' and investors' narrative on the link between interest rate policy and firms' debt maturity choices.[6] Nonetheless, we conclude by bringing the model mechanism to the data. First, we check that large firms increase the frequency of issuance of LT bonds when the policy rate decreases (relatively to small ones). The impact is large: a 25 b.p. decline in EFFR implies an additional 28 b.p. jump for large companies in the probability of issuing new bonds at impact, i.e., 4% of the average likelihood of issuing bonds. Moreover, on the intensive margin, the coupon rate at issuance declines substantially more for large companies. Hence, following an interest rate cut, large firms issue more LT bonds and at lower rates, suggesting that their (relative) reaction is demand driven. Finally, we also test that these buying pressures are associated to portfolio rebalancing by yield-seeking investors. We split corporate bond mutual funds (CBMF) into investment-grade (IG) and high-yield (HY); following Choi & Kronlund (2018), the latter group follows a more yield-oriented investment strategy. In line with the theory, we find a bigger increase in corporate bonds holdings and in portfolio's average maturity for HY-funds after a policy rate decline. The usual 25 b.p. decrease in EFFR prompts a 6 p.p. marginal jump in corporate bonds' holdings for HY-funds and a lengthening in the maturity of held debt-securities by 2 p.p..

CONTRIBUTION TO THE LITERATURE

Our paper contributes to several strands of literature. To start with, our study is the first - up to our knowledge - to provide a systematic analysis of the relation between monetary policy and the

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[5] Importantly, our model leaves room for a standard "balance-sheet channel" of monetary policy, whereby smaller firms benefit more from a relaxation of the monetary conditions. In fact, an interest rate cut relaxes moral-hazard frictions as the value of collateral goes up. Eventually, we prove the existence of equilibria where yield-seeking motives dominate the balance-sheet channel. In this context, the adjustment of the share of LT debt is carried out by the large, unconstrained companies, in line with our evidence.

[6] For instance, a recent article from the Financial Times - commenting the ultra-low interest rate environment during the Covid-19 pandemic - reports that "companies across the US are taking advantage of low borrowing costs to extend the maturity of their debt, selling longer and longer dated bonds to investors starved of yield. (...) As yields have tumbled and investor appetite for debt has remained unsated, corporate treasurers are now making more opportunistic moves." The article also features related comments from CFO of large corporations such as AT&T and from portfolio managers at different investment firms. The article is available at the link: shorturl.at/sHV35.

maturity structure of corporate debt.[7] Few other papers exploit the ex-ante heterogeneity in firms' debt maturity to explain the heterogeneous real effects of monetary policy shocks across firms (Ippolito et al. 2018, Jungherr & Schott 2020b). However, they do not endogenize the response of debt maturity itself to the monetary policy shocks. We do not only provide empirical evidence that such endogenous response is economically meaningful, but also derive a model highlighting a mechanism (operating through short-termist, yield-oriented investors), which can eventually be tested in the data. In a closely related paper to ours, Foley-Fisher et al. (2016) find that a specific unconventional policy by the FED, namely the rebalancing of its portfolio towards longer-term Treasuries, implied a lengthening of the maturity of bond issuance by companies through a gap-filling mechanism (Greenwood et al. 2010). We differ in two dimensions: first, we look at conventional and regular interest rate shocks rather than at a one-time shock to the maturity profile of the FED's portfolio; second, we lever a different mechanism related to yield-oriented investors.

By doing so, we connect to a growing literature stressing the importance of reach-for-yield in financial markets (Becker & Ivashina 2015) for the reallocation of investment across securities in reaction to monetary policy shocks.8 We lever a mechanism, which, following Hanson & Stein (2015), sees yield-seeking investors tilting their portfolios towards LT securities after an interest rate loosening (a mechanism validated empirically with evidence on corporate bonds mutual funds, shown to reach-for-yield by Choi & Kronlund 2018). Our contribution to this literature is to link yield-oriented investors' reaction to monetary policy to the issuance of LT bonds and, ultimately, to the evolution of the maturity structure of corporate debt.

Few novel papers look at financial channels for monetary policy different from a standard credit channel (Bernanke & Gertler 1995), i.e., other mechanisms than bank intermediation. Among others, Foley-Fisher et al. (2016), Grosse-Rueschkamp et al. (2019), Giambona et al. (2020) investigate adjustments in the bond market in response to unconventional monetary policy. Similarly to us, Darmouni et al. (2020) exploit frictions in the Eurozone bond markets and find that access to the bond market is linked to greater firms' sensitivity to interest rate shocks. Ottonello & Winberry (forthcoming) and Jeenas (2018) highlight the importance of default risk

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[7] Gomes et al. (2016) show that nominal long-term debt can generate persistent responses to unanticipated inflation changes (linked to monetary policy shocks) through a debt-overhang mechanism. However, their model does not feature an endogenous firms' optimal debt maturity structure.

[8] For instance, Di Maggio & Kacperczyk (2017) demonstrate that the FED zero rate policy prompted exit in the money-market funds industry and greater risk-taking by surviving funds, with implications for firms borrowing from such institutions. Bubeck et al. (forthcoming) find that negative rates in the Euro Area are associated to enhanced risk-taking in banks' securities portfolio, whereas Lian et al. (2019) and Daniel et al. (forthcoming) obtain homologous findings for individual investors.

and liquid assets for the transmission of interest rate shocks to firms. We innovate by focusing on debt maturity and highlighting a bond-channel of conventional interest rate policy connected to reach-for-yield in financial markets.

We contribute as well to the literature on the determinants of the maturity structure of corporate debt (Barclay & Smith Jr 1995, Berger et al. 2005, Faulkender 2005, Greenwood et al. 2010, Badoer & James 2016), generally placing little emphasis on the policy rate but rather focusing on the term-spread.9 On the other hand, those works discuss extensively the role of financing frictions, but no evidence exists on their interaction with the prevailing monetary policy stance.10

The rest of the paper is organized as follows. In Section 2, we describe the data. In Section 3, we present the baseline empirical findings. To explain them, we elaborate a model in Section 4, which mechanism is tested in Section 5. Section 6 briefly concludes.

1.2 Data

Our empirical analysis covers the period from 1990Q1 to 2016Q4. We employ several datasets, that we describe separately in this section according to the unit-level of analysis.

1.2.1 Time-Series Data

For the time-series analysis of the LT debt share, we use the Federal Reserve Flow of Funds (FoF), tracking financial flows throughout the U.S. economy. We use quarterly data from the credit market liabilities of the non-farm, non-financial, corporate business sector. Our focus rests on the share of total corporate debt with maturity above 1-year, which we label as the share of LT-debt. Following Greenwood et al. (2010), we define short-term debt as the sum of commercial paper and loans with no longer maturity than 1 year (proxied by adding up the FoF entries "other loans and advances" and "bank loans not elsewhere classified"). On the other hand, long-term debt is given by the sum of corporate bonds, mortgages and industrial revenue bonds. The resulting series, given by the fraction of LT debt over the sum of short and LT debt, is depicted in Figure 1. The black line, referring to such variable in levels, displays a generally increasing trend over the period of interest, with the share of LT-debt increasing from roughly 55% to 70%. Throughout the paper, we look at the impact of shocks to the policy rate on the dynamics of the

[9] A notable exception is Baker et al. (2003), showing that the real short-term rate, among other variables, covaries with the corporate LT-debt share.

[10] Relatedly, Poeschl (2017), Xu (2018) and Mian & Santos (2018) ask how cyclical factors impact debt refinancing policy and maturity: our paper differs as it looks specifically at monetary policy.

LT debt share, thereby computing its growth rate over different horizons. In Figure 1, the solid grey line shows the evolution of the quarterly growth rate, that is labelled as ΔLT$-$Debt$_t$ in Table 1 and which mean equals 0.15 p.p.. We also report summary statistics for the cumulative growth rate of the LT debt share over longer horizons, used for pinning down impulse response functions through local projections. In general, the variable ΔLT$-$Debt$_{t+j}$ is computed as the difference between the LT debt share as of year-quarter $t + j$ and $t - 1$, for $j = 0, 1, 2, ..., 20$. For brevity, we show summary statistics only for up to 1-year growth of the LT-debt share. Evidently, both the mean and the volatility of the growth rate increase along with the length of the horizon over which they are computed.

Our first proxy of changes in the policy rate is the quarterly variation in the effective federal funds rate (EFFR), ΔEFFR$_t$, gathered from FRED. Clearly, ΔEFFR$_t$ reflects the evolution of business cycle conditions and the connected endogenous response from the FED. Nonetheless, changes in the EFFR have a "real-world" influence on the firms financing costs. Hence, we first show baseline results based on such raw proxy and next verify the robustness of our findings to employing alternative conventional proxies for exogenous monetary (interest rate) policy shocks. In detail, we borrow data from Gürkaynak et al. (2005), who builds a popular measure of interest rate surprises based on the % change in FED Funds Futures rate in 30minute windows around the policy announcement. Next, we additionally retrieve the Jarocinski & Karadi (2020)'s series of "pure" interest rate surprises, i.e. taking out an informational component - attributed to the provision of private FED information on the state of the economy to private agents through the policy announcement - from the simple variation in the FED Funds Futures rate.

We depict the three series in Figure 2. The post-2009 period is characterized by lower variation in the interest rate policy, associated to the implementation of a zero rate by the FED in the aftermath of the Great Financial Crisis. For this reason, whenever possible, we check that our results survive if we exclude the period from 2009 onward. Moreover, while the three series display a large extent of correlation, both the exogenous variables are in general an order of magnitude smaller than ΔEFFR$_t$. Indeed, while a 1 s.d. change in ΔEFFR$_t$ equals 45 basis points (b.p.), a 1 s.d. change in the Gürkaynak et al. (2005) and Jarocinski & Karadi (2020) shocks amounts to 10 and 8 b.p., respectively (see Table 1).

Finally, we also collect several other FRED macro-economic indicators that we use as controls. In particular, the average annual GDP growth and inflation rates equal 2.46 p.p. and 2.5 p.p., respectively. Moreover, 10% of the year-quarters in our sample are characterized by a recession, as signaled by the dummy Rec$_{t-1}$. The mean quarterly growth rate for the term-spread and the corporate spread are smaller than 1 b.p., though both variables display a large extent of variability (their s.d. amount to 49 b.p. and 25 b.p., respectively).[11] From Thomson Reuters Datastream, we also download information on the share of Treasuries with maturity above 20years and compute its quarterly growth rate (ΔLT−Treas$_{t-1}$).[12]

1.2.2 Firm-level Data

Our primary source for firm-level data is Compustat, a well-known database comprehending balance sheet information on the universe of US listed companies. Using Compustat entails pros and cons. On the positive side, it provides balance sheet information on a quarterly basis, whereas most of other large firm-level datasets contain annual balance sheet only. The relatively higher frequency is desirable in that it aligns better to the frequency of the monetary policy revisions by the FED. On the other hand, the information on debt is rather limited. In fact, we can only distinguish the fraction of total debt with maturity above 1-year - in line with our macroeconomic data from FoF - without further data neither on the maturity profile of existing liabilities nor on the relative weight of bank vs bond financing.

Our sample includes 12,655 companies. Once again, we are mainly interested in the variation over time of the share of debt with maturity above 1-year. The variable ΔLT−Debt$_{f,t+j}$ indeed represents the variation in firm f's LT Debt share from year-quarter $t - 1$ to $t + j$. In Table 1 we report summary statistics for $j = 0, 1, ..., 4$. Across the different horizons, the distribution is centered around 0, as suggested by the median value. Nonetheless, the extent of heterogeneity is remarkable (see the high s.d., increasing along the number of quarters of computation of the cumulative growth rate).

An important variable throughout our analysis is firm's asset size, our preferred proxy for financing constraints, i.e. of access to bond financing. There are large differences in firms' asset size (expressed in logs of 1990q1 millions of US$). From the unconditional summary statistics in Table 1, a one interquartile variation reflect an increase in asset size by nearly 358 p.p.. Clearly, this figure mixes up both cross-sectional and time-series variation.

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[11] The term-spread is defined as the difference between the yield on the 10-year and 3-month benchmark US sovereign bond. The corporate spread reflects risk premium in the corporate sector and is computed as the difference between the Moody's BAA and AAA Seasoned Corporate Bond Yield.

[12] According to the gap-filling theory (Greenwood et al. 2010), the share of LT-debt issued by the corporate sector depends negatively on the share of LT-debt issued by the government. Moreover, Badoer & James (2016) show that corporate debt issuance is especially sensitive to variations in very long-term Treasuries issuance. Hence, controlling for changes in the share of government debt with maturity above 20 years should alleviate concerns that our results are driven by a gap-filling mechanism driven by government debt.

However, our interest in asset-size is aimed at understanding the distribution of the relation between interest rate shocks and maturity structure in the cross-section of firms. To this end, we look at the within (3-digit SIC) industry time-varying distribution of total asset size and define a dummy variable, $\text{Large}_{f,t-1}$, with value 1 if a company is in the upper quartile and 0 otherwise. The choice is due to the fact that - as we will later - it is within this class of firms that debt maturity structure responds to changes in the FED interest rate policy. Moreover, Figure 3 shows how, to start with, the LT debt share is unevenly distributed between firms and increasing in firm size. As a matter of fact, for firms in the first asset-size quartile, the average LT debt share equals roughly 50%, whereas it amounts to nearly 80% for companies in the upper quartiles. Throughout the rest of the paper, we refer to companies in the top-size quartile of their industry distribution as to "large" companies. We also gather additional information from Compustat on other firm level controls such as leverage, liquid assets and sales growth.

To test our mechanism, we then retrieve data on the issuance of bonds with maturity above 1 year from Mergent FISD. Information from Compustat and Mergent FISD are matched through the (6-digit) issuer CUSIP,[13] resulting in a sample of 2,858 bond issuers in Mergent FISD.[14] We start by analyzing bond issuance on the extensive margin through the dummy $1(\text{Issue})_{f,t+j}$, with value 1 if a firm f issues bonds in year-quarter $t + j$ and 0 otherwise,[15] $j = 0, 1, ..., 20$. On average, the likelihood of a current year-quarter new issuance is 6,77%, suggesting that bond issuance is relatively lumpy and infrequent. Such average increases slightly but steadily over future horizons, reflecting the fact that older and/or larger companies tend to issue bonds relatively more frequently. Indeed, Figure 4 looks at the number of bond issuances per year-quarter and splits them depending on whether they are conducted by a large company or not. The share of new issuances by large companies is disproportionately large. In fact, while such firms account (by construction) for roughly 1/4 of the firms, they represent about 60% of new bond issuances. In other terms, this is prima-facie evidence that large companies are much more active in the corporate bond market and hence more likely to react to potential variation in the associated

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[13] In a couple of dozen of cases, multiple companies - typically 2 or 3 - have the same 6-digit CUSIP in Compustat, referring to different subsidiaries of a same group. In such cases, we retain the largest company in Compustat among the ones with same 6-digit CUSIP in an effort to identify the mother company. Excluding all such companies would not affect the results.

[14] Such number of firms refers to the companies in our regression sample. Since we apply firm fixed effects in our regressions, those are companies that issue bonds at least twice throughout our period of analysis.

[15] Assigning 0 to periods in which Mergent FISD does not report a company's bond issuance requires to know whether a company is active or not. To this end, we label a company as "active" if it reports balance sheet information in Compustat, and as "inactive" if it does not. In practice, disappearance from Compustat means that a firm has delisted, implying that it cannot issue bonds anymore.

financing costs. In particular, as we are interested in the response of financing costs to monetary policy, we retain data on the annualized coupon rate at issuance, $\text{CouponRate}_{f,t+j}$, which is equal to 6% on average.[16]

1.2.3 Corporate Bonds Mutual Funds Data

We retrieve data on corporate bond mutual funds (CBMF)[17] holdings from the CRSP

Survivor Bias-Free dataset, including information on both surviving and dead funds. Following Choi & Kronlund (2018), we split funds into High Yield (HY) and Investment Grade (IG) funds based on standard Lipper style codes.[18] We label HY-funds as yield-oriented: as shown by Choi & Kronlund (2018), they intuitively invest relatively more in longer and riskier debt-securities.[19] Ideally, one would build a measure of fund-specific reach-for-yield, but this requires security-level data on CBMFs' holdings that we do not have access to. Hence, we use the just described secondbest, empirically grounded HY-vs-IG funds proxy.

Overall, we analyze 3,487 funds (2,034 are IG and 1,453 are HY) over the dataperiod 2010q2-2018q2. Table 1 describes the related summary statistics. A first outcome variable of interest is the cumulative growth rate of corporate bond holdings through time, $\Delta\text{CB}_{m,t+j}$, which displays a large extent of heterogeneity across funds. Second, we are also interested in the changes in the fund's average (weighted) portfolio maturity over time, $\Delta\text{Matu}_{m,t+j}$, equally showing significant differences in the cross-section of funds. We gather additional information on fund characteristics, used as controls in our models, including the fund's turnover and expenses ratio, the net asset value and returns.

1.3 Empirical Analysis

In this section, we present the baseline empirical findings of our paper. First, we present the aggregate-level analysis. Next, we investigate cross-sectional differences across firms.

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[16] The large fall in observations with respect to those for the variable $1(\text{Issue})_{f,t+j}$ reflects the fact that the distribution of $Coupon_{f,t+j}$ is conditional on $1(\text{Issue})_{f,t+j} = 1$.

[17] Specifically, like Choi & Kronlund (2018), we limit the sample of funds to CRSP style categories I, ICQH, ICQM, ICQY, ICDI, ICDS, or IC.

[18] IG funds are classified as those with a Lipper style code of either A, BBB, IID, SII, SID, or USO and HY funds are those coded HY, GB, FLX, MSI, or SFI.

[19] We refer to the data in Table 1, Panel B from Choi & Kronlund (2018) and to the related discussion in the paper. Interestingly, HY funds reach-for-yield relatively more than IG funds, i.e. they invest in securities with higher yields. Importantly, this difference is explained by both risk and maturity, whereas IG funds tend to reach-for-yield more within a given bucket of risk-maturity bucket.

1.3.1 Time-series Analysis

We apply local projections (Jordà 2005) to study the response of the share of LT-debt to changes in the FED's interest rate policy. In particular, we estimate separately the following regressions through OLS:

$$\Delta_h y_{t+h} = \beta_{1,h} \Delta EFFR_t + \Gamma_h X_{t-1} + u_{t,h} \qquad (1)$$

for h = 0, 1, ..., 20. The dependent variable, $\Delta_h y_{t+h}$, is given by the cumulative variation in the share of LT-debt between year-quarters $t - 1$ and $t + h$. Hence, plotting the coefficients $\beta_{1,h}$ provides the impulse-response function of the share of LT debt to a change in the EFFR as of year-quarter t, $\Delta EFFR_t$. Moreover, X_{t-1} is a vector of lagged macro-controls, including variables which might simultaneously have an influence on $\Delta_h y_{t+h}$ and on the current interest rate policy. In particular, X_{t-1} comprehends: the annual GDP growth rate and inflation rate; the quarterly variation in the 10y-3m term-spread, in the corporate spread and in the share of Treasuries with maturity above 20-year; a recession dummy. Finally, $u_{t,h}$ is a robust error-term.

Figure 5 reports the impulse-response function obtained from the OLS-estimation of the coefficients $\beta_{1,h}$ in Equation 1. In particular, the plot assumes a 25 b.p. quarterly negative variation in the EFFR - i.e., a loosening of the short-term policy rate, a convention we maintain throughout the rest of the paper - and also displays the 10% confidence interval around the point estimates. Clearly, an interest rate cut boosts the share of LT-debt. The effect is very persistent and, while effective at impact, peaks up 3 years after the shock, and does not fade away throughout the considered 5-year time-window. Such implausibly large degree of persistence might reflect the endogeneity of the simple raw variation in the EFFR and its significant autocorrelation along monetary policy cycles. That said, the effect is economically meaningful. For instance, a 25 b.p. interest rate descent in year-quarter t implies a cumulative increase in the share of LT-debt by 0.42 p.p. one year after. For comparison, the average growth rate of the LT-debt share equals 0.15 p.p. on a quarterly basis, and 0.83 p.p. on an annual basis.

Importantly, we validate that such result is robust to employing alternative and exogenous monetary policy shocks. That is, in Equation 1, we replace $\Delta EFFR_t$ with $\varepsilon_t^{mp,g}$ $\varepsilon_t^{mp,jk}$, i.e., the high-frequency surprises on FED Funds Futures rates from Gürkaynak et al. (2005) and the related pure interest rate shocks from Jarocinski & Karadi (2020), respectively. Figure 6 shows the resulting impulse-response functions, calibrated for a 1 s.d. expansionary exogenous shock. The analysis validates the positive effect of an interest rate loosening on the LT-debt share.

Quantitatively speaking, the influence of both shocks is similar, and aligned to that of a 25 b.p. reduction in the raw EFFR. Nonetheless, the shock displays a much less persistent effect, with the impact on the share of LT-debt vanishing in 5 or 14 year-quarters when using the Jarocinski & Karadi (2020) and Gürkaynak et al. (2005) shocks, respectively.

Finally, we also verify in the Empirical Appendix Figures A1 and A2 that the findings survive to restricting the sample to the period between 1990q1 and 2008q4, which we label as pre-crisis. This is an important robustness check in that most of the variation in the FED's interest rate policy occurs before 2008.

1.3.2 Firm-level Analysis: Econometric Model

We investigate US listed firms' quarterly balance-sheets in order to understand crosssectional differences in the response of the LT-debt share to variations in the interest rate policy. To this end, we borrow the empirical strategy from Jeenas (2018) and Ottonello & Winberry (forthcoming), using a panel version of the Jordà (2005)'s local projections. In practical terms, we estimate by OLS the following set of equations:

$$\Delta_h y_{f,t+h} = \beta_{1,h}\Delta EFFR_t + \beta_{2,h}Large_{f,t-1} + \beta_{3,h}Large_{f,t-1} * \Delta EFFR_t + + \Gamma_h X_{f,t-1} + \mu_f + \mu_{s,t} + u_{f,t+h} \quad (2)$$

for $h = 0, 1, ..., 20$. The dependent variable, $\Delta_h y_{f,t+h}$, is given by the cumulative variation of the share of LT-debt of firm f between year-quarters $t - 1$ and $t + h$. Most importantly, the model includes the full interaction of the raw quarterly change in EFFR, $\Delta EFFR_t$, and a dummy for large companies, i.e., with value 1 for companies in the upper quartile of the respective industry asset-size distribution, $Large_{f,t-1}$. The coefficient of main interest is $\beta_{3,h}$, capturing the relative response of large companies (as compared to smaller ones) to a variation in the FED short-term policy rate.

We augment the model with a vector of firm controls, which comprehends the (lagged) share of liquid assets, leverage and sales quarterly growth. Eventually, such variables are also fully interacted with $\Delta EFFR_t$. By doing so, we horse-race our channel (based on firms size as a proxy for bond financing constraints) against other layers of heterogeneity which have been found to influence firms' response to monetary policy shocks.[20] Furthermore, we interact $Large_{f,t-1}$ with the usual set of macrocontrols for avoiding that $\beta_{3,h}$ reflects contemporaneous response of large

[20] Jeenas (2018) shows that companies with a relatively lower share of liquid assets respond more to monetary policy shocks. Ottonello & Winberry (forthcoming) find that distance to default matters as well for firms' reaction to monetary policy, with leverage being a good proxy for it. Sales growth is meant to capture a firm's profitability.

companies to other shocks, which may correlate with the FED's interest rate policy decisions. The model is saturated with firm and indyustry*year-quarter fixed effects, i.e. μ_f and $\mu_{s,t}$, respectively. The former set of dummies controls for all observed and unobserved time-invariant heterogeneity at the level of the firm; the latter absorbs time-varying shocks which are common to firms in a given (3-digit SIC) industry. The application of such fixed effects implies that our coefficient of interest $\beta_{3,h}$ is identified by: i) within-firm variation over time, i.e., changes in response of LT-debt share by an otherwise identical firm when it is large as compared to when it was small; ii) cross-sectional variation across firms in a given industry. Finally, $u_{f,t+h}$ is an error term, which we double-cluster at the firm and industry*yearquarter level.

However, the relative adjustment of large companies estimated through Equation 2 does not allow to understand the overall response of both large and smaller companies. In fact, Equation 2 is saturated with industry*year-quarter fixed effects, which span out completely time-series variation common across all firms.

Hence, we additionally estimate the following model separately for firms in different size-quartiles:

$$\Delta_h y_{f,t+h} = \beta_{1,h}\Delta EFFR_t + \Psi_h X_{f,t-1} + \mu_f + \nu_{f,t+h} \qquad (3)$$

That is, we estimate a model which exploits just time-variation and hence describes the absolute change in the share of LT-debt after a change in interest rate by the FED.[21] In fact, we do not use year-quarter fixed effects (nor any subtler version of them), while we keep using firm fixed effects to control for unobserved timeinvariant firm heterogeneity.

The trade-off across the two models is clear: Equation 2 precisely estimates the cross-sectional differences across firms, as it controls for time-varying common heterogeneity within narrowly defined industries. On the other hand, model 3 pins down the absolute variation in LT-debt share associated to variation in the policy rate. Hence, it serves the purpose of better understanding the connection between firm-level and time-series findings.

1.3.3 Firm-level Analysis: Results

Figure 7 plots the impulse-response function obtained from the estimation of the parameters $\beta_{3,h}$ - for h = 0, 1, ..., 20 - from Equation 2. Relatively to smaller firms, large companies expand LT-

[21] In Equation 3, $X_{f,t-1}$ is the usual vector of macro and firm-level controls.

debt more when the policy rate goes down. That is, large companies react more in line with the aggregate-level evidence shown in section 3.1. For understanding the absolute response-level, however, we additionally estimate Equation 3 within different size-quartiles. The resulting impulse response functions are displayed in Figure 9 and suggest that only large companies do adjust, whereas smaller firms' LT-debt share is generally insensitive to monetary policy. Moreover, we replicate both exercises using the by-now familiar exogenous monetary policy shocks from Gürkaynak et al. (2005) and Jarocinski & Karadi (2020), obtaining specular findings (see Figures 8 and 10).

For gauging the economic significance of the just commented effects, we refer to the baseline figures employing the endogenous quarterly change in the EFFR. Once again, Figures 7 and 9 are calibrated to a 25 b.p. (expansionary) negative EFFR variation. The jump in the share of LT-debt by large companies peaks up 6 yearquarters after the policy change, when it is comprised between 0.45 and 0.55 p.p. (depending on whether one takes as a reference the adjustment in Figure 9 or in Figure 7, respectively).[22] Interestingly, the described size of the effect is comparable at relevant horizons - with that observed at the aggregate level. At the firm-level, the 6 year-quarter cumulative growth of the LT-debt share equals -1.53 p.p. on average (not shown for brevity in Table 1).

Size turns out being the key firm-level attribute to explain cross-sectional differences across firms. In this respect, we report in Table 2 additional coefficients from the estimation of the baseline firm-level model (Equation 2).[23] In particular, we show the horse-race with the other balance-sheet characteristics employed as firm-level controls. First, companies tend to increase the share of LT-debt when sales jump; nonetheless, the interaction of such dynamics with monetary policy is insignificant. Moreover, intuitively, the share of LT-debt goes down when firms hold relatively more liquid assets, reflecting maturity matching of assets and liabilities. Also in this case, however, the share of held liquid assets does not influence the relation of debt maturity structure and monetary policy. Similarly, leverage has a small and marginally significant influence on such relation at impact, but the effect fades away already 1 year-quarter after, whereas at the macroeconomic level the policy rate has a more persistent impact on the share of LT-debt.

[22] From a formal perspective, the absolute variation in LT-debt share is pinned down in Figure 9. Nonetheless, we also refer to Figure 7 as it estimates precisely the relative adjustment of large companies and the baseline effect on smaller ones can be placed at 0.
[23] Table 2 displays coefficients associated to a 1 p.p. (contractionary) increase in the EFFR.

Differently, the effect of our dummy for large companies is strongly significant and persistent over time, and resembles well the patterns observed at the aggregate level.[24]

Finally, for robustness, we repeat the exercise over the pre-crisis period (i.e., from 1990 to 2008, included) so to restrict our analysis to a time-window with substantial interest rate shocks. Results are reported in Appendix Figures A3 and A4 and confirm the baseline findings both qualitatively and quantitatively.

1.4 Model

In this section, we present a model which explains our empirical findings: i) the aggregate-level share of LT-debt goes up following an interest rate loosening; ii) such effect is entirely driven by the adjustment of very large companies. Existing theoretical frameworks are not useful in this respect. In fact, models of monetary transmission to firms do not include an explicit discussion on debt maturity, whereas the corporate finance literature does not typically focus on the policy rate.[25] First, we present the model setup. Next, we characterize the equilibrium conditions and perform comparative statics exercises which pin down the relation between the short-term policy rate and the firms' debt maturity structure. Finally, we perform few empirical test to validate the mechanism proposed by the model.

1.4.1 Setup

Our economy lasts three periods ($t = 0, 1, 2$) and is populated by a continuum of firms and investors. Each firm is endowed with capital A - a proxy for firm size - and a project. A is heterogeneous across firms and distributed uniformly across firms on the interval $[0, I]$, where I is the initial investment into each firm's project in period 0. Each project also features a stochastic re-investment ρ in period 1, drawn out of an exponential distribution $f(\rho) = {}_{\chi}e^{-\chi\rho}$, for $\rho \in [0, \infty)$.

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[24] The coefficient on the large company dummy alone turns out being insignificant. This result apparently in contradiction to the stark differences in the share of LT-debt observed across size quartiles in Figure 3 - is due to the inclusion of firm fixed effects in our regressions. Hence, accordingly with our aim of exploring cross-sectional differences, the relevant variation in size captured by our dummy operates mostly between companies, rather than within.

[25] In corporate finance, the focus rather rests on the term-spread, as it describes the relative cost of long-term debt relative to short-term debt. In particular, a jump in the term-spread would predict an increase in the relative cost of LT debt, and hence a related decrease in the issuance of LT bonds - a prediction contradicting our findings. As a matter of fact, a policy rate loosening predicts an expansion of the term spread (see e.g. Adrian & Shin 2010), a robust relation which we document to hold in our sample. In Appendix Figure A5, a simple scatterplot - with the quarterly change in the EFFR on the x-axis and the variations of of the term-spread on the y-axis - suggests a strong negative relation. We test such influence more formally in Appendix Figures A6 and A7, which reports the impulse-response function from a model with the term-spread as dependent variable but otherwise identical to that in Equation 1.

If the reinvestment need is not met, the project is liquidated and does not generate any payoff in period 2.

Each project generates a riskless short-term pay-out r in period 1. Differently, in period 2, conditional on the reinvestment need being satisfied, the project yields R in case of success, and zero in case of failure. The likelihood of success depends on firms' behavior. We assume that if a firm exerts effort, the project is successful with probability p_h (without loss of generality, we set $p_h = 1$); on the other hand, if the firm shirks, success materializes with probability $p_l < 1$, but it enjoys private benefits B. Additionally, there is aggregate risk: with probability $1 - \delta$, all firms get a pay-out of zero in period 2. Firms do not have a storage technology and are protected by limited liability.

Investors are competitive and a subset of them features reach-for-yield behavior, meaning that they care about the return on their portfolios relative to the current interest rate, as opposed to the series of current and future interest rates. They therefore take more risk than rational investors if interest rates are low.

The short term interest rates between period 0 and 1, and 1 and 2, respectively, are set exogenously by a monetary authority. These exogenous short-term interest rates are denoted as i_1 for the interest rate from period 0 to period 1, and as i_2 for the interest rate from period 1 to period 2. All agents in this model use these short-term rates as discount rates and we treat i_1 as the policy rate. Investors can lend directly at these exogenous rates, while firms receive funds intermediated by the investors.

In the following, we lay out formally how firms and investors are modeled.

Firms. We follow Holmström & Tirole (1998, 2000) in modeling firm financing as a moral hazard problem in which each individual firm is an agent.

We assume that firms can credibly commit to a contract stating that the project is carried on to period 2 whenever the stochastic re-investment is sufficiently small, i.e. if $\rho \leq \rho^*$, and terminated otherwise. The continuation threshold ρ^* is a choice variable of the firm and is common knowledge. In equilibrium firms will differ in their choice of ρ^*, thus we use the notation $\rho^*(A)$, denoting the choice of a firm with capital A.

To finance the gap between the initial investment and the endowment, firms issue short and long-term debt. A riskless short-term bond is sold at price $P_s = 1/(1 + i_1)$ at time 0 and is promised to yield 1 in expectation in period 1, while a long-term bond is sold at price P_l in time 0 and yields 1 in period 2 if the project is successful. The amount of issued short-term and long-term debt is denoted by d_s and d_l, respectively.

The timing of the financing and execution of the project is as follows. First, in period 1, short-term creditors are compensated out of earnings r, as the firm must assure to repay d_s. Next, the firm draws the re-financing shock from $f(\rho)$. If the decision is not to refinance - i.e., if $\rho > \rho^*$ - then the firm abandons the project and consumes what is left, whereas long term bond-holders do not receive any compensation. If the project is continued - i.e., if $\rho \leq \rho^*$ - and turns out to be successful in period 2, the entrepreneur enjoys $R_b = R - d_l$, while long-term bond-holders receive their compensation d_l. Eventually, if the project is unsuccessful then the firm is again liquidated at value zero and neither the long-term bond-holders, nor the entrepreneur, receive anything. Hence, it follows that, to induce the entrepreneur to exert effort, the following condition must hold: $R-\frac{B}{\Delta p}\geq d_l$.

where $\Delta p = p_h - p_l$. Intuitively, this incentive compatibility constraint means that the repayment in an optimal contract cannot be too large, otherwise the entrepreneur will shirk.

Moreover, limited liability and riskless short-term debt implies: $\rho^* \leq r-d_s$, i.e., the firm cannot be asked to meet the liquidity shock with other funds than those stemming from the project returns. Finally, the firm must raise (and investors must be willing to provide) enough money to finance the project in the first place: $\frac{1}{1+i_1}d_s+P_l d_l\geq I-A$.

Taking stock, the general problem of the firm reads (from now on we will index the choice variable by the endowment):

$$\max_{\rho^*(A),\, d_s(A), d_l(A)} \frac{r}{1+i_1}-\frac{\int_0^{\rho^*(A)}\rho f(\rho)d\rho}{1+i_1}+\frac{\delta F\left(\rho^*(A)\right)}{(1+i_1)(1+i_2)}R+\left(P_l(A)-\frac{\delta F\left(\rho^*(A)\right)}{(1+i_1)(1+i_2)}\right)d_l(A) \qquad (4)$$

subject to:

$$r-\rho^*(A)\geq d_s(A) \quad \text{(LL)}$$

$$R-\frac{B}{\Delta p}\geq d_l \quad \text{(IC)}$$

$$\frac{1}{1+i_1}d_s+P_l d_l\geq I-A \quad \text{(IR)}$$

The objective function represents expected firm profits, discounted as of t = 0. The first term gives the risk-free period-1 revenues and the second one subtracts the expected period-1 payments due to the liquidity shock. The third element provides the expected period-2 revenues, influenced by both idiosyncratic liquidity shock and aggregate risk. The last term of the equation collects the net proceeds from the issuance of LT bonds, i.e., the value of liquidity minus total expected repayments.

Investors. Firms borrow in bond markets featuring a continuum of investors. Investors are heterogeneous and we denote their type as j. Investors have zero initial wealth and construct long-short positions to maximize:

$$E[w_j]-\frac{\gamma}{2}Var[w_j]$$

where w^j is wealth of an investor of type j as of t = 2. They purchase a portfolio of LT debt, issued by the firms, and finance this position by rolling over short-term borrowing. As a result, w^j equals:

$$w_j = d_l^{*j} - \iota(i_1, i_2, j) \int_0^I P_l(A) d_l^j(A) dA.$$[26]

Here, d_l is the realized payoff from holding a portfolio comprising LT debt of all firms, whereas $\iota(i_1, i_2, j)$ is the individual (compound) factor that each investor uses to judge her financing costs. $\iota(i_1, i_2, j)$ is heterogeneous across investor types. In particular, we assume two investor types: $j \in \{R,Y\}$. A fraction $1 - \alpha$ of the investors are "rational" and their compounded discount rate $\iota(i_1, i_2, R)$ is $(1 + i_1)(1 + i_2)$. On the other hand, a fraction α of the investors is of the "yield-seeking" type, whose modelling we borrow from Hanson & Stein (2015). Specifically, such yield-seeking investors compare the expected returns from their investments only with the current interest rate instead of the stream of expected interest rates. Their $\iota(i_1, i_2, Y)$ is $(1 + i_1)^2$.[27] The explicit expectation and variance of investor wealth at t = 2 are:

$$E[w^j] = \delta\int_0^I F(\rho^*(A))\, d_l^j(A)dA - \iota(i_1, i_2, j)\int_0^I P_l(A)d_l^j(A)dA$$
$$Var[w^j] = Var[d_l^{*j}] = \left(\int_0^I F(\rho^*(A))\, d_l^j(A)dA\right)^2 \delta(1-\delta).$$

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[26] We assume that the mass of firms is *I*, such that the density of each firm type *A* is 1 and can be omitted from notations.
[27] The fact that these investors are discounting using an incorrect rate generates the yield-seeking behavior. This modeling choice can be justified by agency or accounting considerations that lead investors to worry about short-term measures of reported performance.

In the expression for expected wealth, the expected revenues reflect the fact that, for a firm with endowment A, the likelihood of repayment equals $\delta F(\rho^*(A))$. Next, in the variance term, $\int_A F(\rho^*(A))\, d_l^j(A)$ is treated like a constant due to full diversification of firms' idiosyncratic risk; in other terms, investors' risk only depends on aggregate shocks.

Investors maximize their wealth by optimally choosing a LT debt portfolio including all firms' debt and, due to the mean-variance utility assumption, they have limited risk-bearing capacity. We assume that there two types of investors. The general problem for both types is:

$$\max_{d_l^j(A)} \delta \int_0^I F(\rho^*(A))\, d_l^j(A)dA - \iota(i_1, i_2, j) \int_0^I P_l(A) d_l^j(A)dA - \frac{\gamma}{2}\left(\int_0^I F(\rho^*(A))\, d_l^j(A)dA\right)^2 \delta(1\text{-}\delta) \quad (5)$$

Finally, we assume that there is an inelastic demand g, originating from preferredhabitat investors into LT debt, that is defined in terms of expected bond payments in $t = 2$.[28]

1.4.2 Discussion of the Setup

It is important to discuss the modeling choices and assumptions we made in our model. The model by Holmström & Tirole (1998, 2000) provides a tractable framework to study how the maturity structure interacts with financing constraints and investors demand. In their model, the defining difference between short-term and long-term financing is credit risk that affects only debt of longer maturity. As Holmström & Tirole (1998, 2000), we abstract from other sources of risk, such as duration or rollover risk. Our goal is to endogenize the maturity structure of liabilities of the corporate sector to compare theoretical predictions with the empirical regularities discussed above. The key model ingredients for this are the intermediate income r, which provides firms a cash-flow to service short-term debt, and the incentive compatibility constraint, which restricts firms long-term debt choice. The additional trade-off of more continuation against more short-term debt, adds a margin of adjustment for the maturity structure of constrained firms. Since our investors are risk-averse and thus more complex than those in Holmström & Tirole (1998, 2000), we make some simplifying assumptions, namely that there is no storage technology, and that short-term debt is riskless. The latter can be rationalized when the intermediate cash-flow is large enough, so the firm is sufficiently "cash-rich", then, once its short-term debt has reached $r - \rho^*$ no firm would not want to increase short-term debt above this riskless amount, as the price would

[28] Greenwood et al. (2010) make a similar assumption and describe such investors as pension funds, life insurance companies, endowments, or any institution with an inelastic demand for long-term assets.

deteriorate too fast for an increase in short-term debt to be revenue-increasing. As in Holmström & Tirole (1998, 2000), we abstract from firms rolling short-term debt over to t = 2. This allows us to find a well-defined maturity structure and is without loss of generality: as all contingencies and their probabilities are known by firms and investors in t = 0, and the contract specifies how to deal with them, there will be no incentive to refinance once the liquidity shock has realized.

There has long been an established theoretical literature on firms' optimal debt maturity choice such as Flannery (1986, 1994), Diamond (1991), Diamond & He (2014), He & Milbradt (2016). The downside of the aforementioned papers is that they do not focus on firms' financial constraints. Our empirical evidence suggests that financing constraints are a crucial element of the explanation and thus we use Holmström & Tirole (1998, 2000) as starting point of our analysis. Other papers consider the effect that a given debt maturity has on financial outcomes, such as roll-over risk and credit risk (He & Xiong 2012a,b). Relative to these papers, we are exploring the relationship in the opposite direction, in that we are trying to understand how changes in financing conditions affect maturity choices.

Moreover, as we show below, a model without yield-seeking investors does not match these empirical facts. This is our motivation to extend such that investor demand is sensitive to an interest rate change. We achieve this by introducing riskaversion and reach-for-yield behavior for some investors.[29]

1.4.3 Equilibrium

First we formally characterize the competitive equilibrium of our model.

DEFINITION 1. A competitive equilibrium is a set of quantities $\{d_s(A), d_l(A), d_l^R(A), d_l^Y(A)\}_{A\in[0,I]}$, cut-off rules $\{\rho^*(A)\}_{A\in[0,I]}$ and prices $\{P_l(A)\}_{A\in[0,I]}$ such that:

1. $\{d_s(A), d_l(A), \rho^*(A)\}_{A\in[0,I]}$ solve firms' optimization problem (4), given $\{P_l(A)\}_{A\in[0,I]}$.
2. $\{d_l^R(A), d_l^Y(A)\}_{A\in[0,I]}$ solve rational and yield-seeking investors' respective maximization problems (5).

[29] There are various approaches to model reach-for-yield behavior, such as Drechsler et al. (2018), Acharya & Naqvi (2019), Lu et al. (2019), Campbell & Sigalov (2020). We follow the approach of Hanson & Stein (2015) who model reach-for-yield as a subset of agents using the current interest rate to discount future income, instead of the path of expected future interest rates.

The LT bond market clears:

$$d_l(A) = (1-\alpha)\, d_l^R(A) + \alpha d_l^Y(A) + \frac{g}{\int_0^I F(\rho^*(A))dA}.^{30} \quad (6)$$

We start with the first-order conditions for "rational" and "yield-seeking" investors, respectively:

$$P_l(A) = \frac{\delta F(\rho^*(A)) - \gamma F(\rho^*(A))\delta(1-\delta)\int_0^I F(\rho^*(A))\, d_l^R(A)dA}{(1+i_1)(1+i_2)} \quad (7)$$

$$P_l(A) = \frac{\delta F(\rho^*(A)) - \gamma F(\rho^*(A))\delta(1-\delta)\int_0^I F(\rho^*(A))\, d_l^Y(A)dA}{(1+i_1)^2} \quad (8)$$

where $d_l^Y(A)$ and $d_l^R(A)$ are the demand of firm A's bonds by yield-seeking and rational investors, respectively. The two investor types compete in the same market to buy LT debt and face the same price. However, their demand differs due to a different attitude towards interest rates. The first term in the right-hand side of both expressions gives the expected payoff from holding a firm A's LT bonds. Importantly, for such an expected payoff, yield-seeking investors are willing to pay a premium on LT debt if $i_2 > i_1$, as they overreact to changes in i_1, relative to rational investors. The second term suggests that investors, being risk averse, are compensated (through a lower price) for holding risky LT debt. Furthermore, a smaller continuation cutoff $\rho^*(A)$ implies a lower price, as the probability of repayment in t = 2 descents.

Rearranging equations (7)-(8) and plugging them into the market clearing condition (6) yields the inverse demand for firm A's LT-debt:

$$P_l(A) = \delta F(\rho^*(A))\, \frac{1-\gamma(1-\delta)\left(\int_0^I F(\rho^*(A)d_l(A)dA - g\right)}{(i+i_1)[\alpha(i+i_1)+(1-\alpha)(1+i_2)]} \quad (9)$$

The discount factor in (9) is a weighted average of the discount factors of the two types of investors. Moreover, the willingness to pay of the marginal investor decreases in the aggregate volume of LT debt held. Before proceeding further, we state a lemma that will be useful for the derivations of our key results.

LEMMA 1. The price for LT debt of a firm with endowment A is unaffected by changes in the supply of LT debt $d_l(A)$ and increases in $\rho^*(A)$.

---

[30] The inclusion of a large enough *g* ensures that, under any circumstances, all firms borrow a positive amount of LT-debt.

Proof. See the Theory Appendix. □

Next, the first order conditions of the firms’ maximization program are:

$$\lambda_3(A) = \lambda_1(A) \quad (10)$$

$$P_1(A) = \frac{1+\lambda_2(A)}{1+\lambda_3(A)} \frac{\delta F(\rho^*(A))}{(i+i_1)(1+i_2)} \quad (11)$$

$$f(\rho^*(A)) \left[\frac{\delta R}{1+i_2} - \rho^*(A)\right] + \left((i+i_1)(1+\lambda_3(A)) \frac{\partial P_1(A)}{\partial \rho^*(A)} - \frac{\delta f(\rho^*(A))}{1+i_2}\right) d_1(A) - \lambda_1(A) = 0 \quad (12)$$

where λ_1, λ_2, λ_3 are the multipliers linked to the three constraints LL, IC and IR, respectively. Condition (10) signals that LL binds if and only if IR does. In Equation (11), the firm valuation of one unit of LT debt equals the NPV of the risky project times a factor positively (negatively) related to the tightness of the IC (IR) constraint. Finally, in condition (12), the optimal liquidation cutoff decreases when the LL constraint binds relatively more.

1.4.4 Unconstrained and Constrained Firms

In this section we analyze the conditions for the existence of unconstrained and constrained firms in equilibrium and characterize their optimal plans. Focusing on such equilibrium, rather than others where all firms are either constrained or unconstrained (in the sense specified below), allows us to relate our theory to the crosssectional empirical findings in Section 3.2. We call a firm "unconstrained" whenever all three constrains are slack and "constrained" if the opposite is true.

PROPOSITION 1. The three constraints of the firm problem only bind concurrently. Unconstrained and constrained firms coexist in equilibrium if:

$$\bar{A} \in (0, I)$$

where:

$$\bar{A} = I - \frac{r - \frac{\delta R}{1+i_2}}{1+i_1} - \frac{F\left(\frac{\delta R}{1+i_2}\right)}{(1+i_1)(1+i_2)} \delta \left(R - \frac{B}{\Delta P}\right)$$

Proof. See the Theory Appendix. □

$\overline{A}$ denotes the lowest endowment at which the optimal continuation threshold is just feasible if the firm takes on the largest possible level of (short and long-term) debt. In practice, $A = \overline{A}$ is a threshold value for firm size, above (below) which firms can (cannot) implement the optimal continuation value.

We now proceed with the description of the optimal plans for unconstrained and constrained firms. For unconstrained firms $\rho^*(A)$ is at the optimal value:

$$\rho^*(A)=\frac{\delta R}{1+i_2} \text{ if } A\geq \overline{A} \quad (13)$$

which follows from condition (10) and represents the risky project's revenues discounted of aggregate risk. For these firms, the limited liability constraint does not bind, so they are indifferent with respect to taking any amount of short-term debt. Moreover, they collectively invest into LT debt until its price equals their own valuation, as reported in Equation (11).

We find it useful to make an assumption on how LT ~~debt~~ is distributed between unconstrained firms, so to match the stylized empirical fact in Figure 3, i.e., unconstrained firms have a larger LT debt share than constrained firms. Concretely, we assume that unconstrained firms choose a combination of LT debt and short-term debt allowing them to match the highest LT debt ratio among constrained firms. [31]Differently, a constrained company piles up as much short- and long-term debt as they need, namely for $A < \overline{A}$:

$$d_l(A) = R-\frac{B}{\Delta P}$$

$$d_s(A) = r-\rho^*(A)$$

Finally, the next lemma shows that the continuation cutoff of constrained firms is set below the optimal level.

LEMMA 2. Assuming $\frac{\chi\delta}{1+i_2}\left(R-\frac{B}{\Delta P}\right)<1$, a constrained firm chooses a continuation value $\rho^*(A)\in[0, \frac{\delta R}{1+i_2})$. It also follows that:

[31] Details in the Theory Appendix in the proof of Proposition 2. This assumption is made possible by an appropriately large preferred-habitat investor demand *g*. In fact, it implies a relatively large excess-demand for LT-debt, which is filled by unconstrained companies.

Proof. See the Theory Appendix. □

Intuitively, firms below $\overline{A}$ need a large amount of debt. In particular, high reliance on short-term financing tightens the limited liability constraint, so that ρ^* is compromised. Increased firms own capital (i.e. size A) reduces the need for short-term finance, therefore increasing the chosen continuation cutoff for constrained companies. Similarly, a descent in the short-term rate i_1 boosts the price of short- and LT debt, thereby loosening the IR constraint, which co-moves with the LL constraint by condition (10). A relaxation of the LL constraint increases the feasible continuation value. Clearly, both relations do not apply among unconstrained companies as any firm with $A > \overline{A}$ is already at the optimum.

1.4.5 Effects of Poilicy Rate Change on the Maturity Structure

In this section, we study how changes to the short-term policy rate i_1, controlled by the monetary authority, affect firms' debt maturity structure. We first derive predictions in a setting without yield-oriented investors and next in the baseline model presented above to highlight the fact that the inclusion of reach-for-yield motives allows us to replicate the empirical facts documented above.

EFFECT OF OF POLICY RATE CHANGE WITHOUT YIELD-SEEKING INVESTORS

The absence of reach-for-yield motives represents a limiting case of the above model in which $\alpha = 0$. In such setting, rational investors have no incentive to hold LT debt, because it is risky and firms are just willing to price debt in a risk-neutral fashion.[32] In practical terms, the inverse demand function for LT bonds equals:

$$P_l(A)=\delta F\big(\rho^*(A)\big)\frac{1\text{-}\gamma(1\text{-}\delta)\left(\int_0^I F(\rho^*(A)d_l(A)dA\text{-}g\right)}{(i\text{+}i_1)(i\text{+}i_2)}$$

Market clearing implies:

$$\frac{1\text{-}\gamma(1\text{-}\delta)\left(\int_0^I F(\rho^*(A)d_l(A)dA\text{-}g\right)}{(i\text{+}i_1)(i\text{+}i_2)}=\frac{1}{(i\text{+}i_1)(i\text{+}i_2)}$$

[32] This can be seen in Equation (11). For unconstrained companies, $\lambda_2 = \lambda_3 = 0$, so that the resulting price clearly just discounts aggregate and idiosyncratic risk, without offering any risk-premium.

or:

$$\int_0^{I} P_{l}(A)d_l(A)dA = \int_0^{\overline{A}} F(\rho^*(A)d_l(A)dA + \int_0^{\overline{A}} F(\rho^*(A)d_l(A)dA = g \quad (14)$$

That is, the LT bonds issued by constrained and unconstrained firms have to net out. In this case, in reaction to a descent in the interest rate i_1, by Lemma 2, constrained firms increase their continuation value ρ^*. The right hand side of (14) is constant, and thus unconstrained firms must decrease their LT debt, therefore shortening debt maturity. This result clashes with the cross-sectional empirical evidence.[33]

EFFECT OF POLICY RATE CHANGE WITH YIELD-SEEKING INVESTORS

The next proposition resumes the effects of a policy-rate change in our full model. First, for the sake of exposition we introduce a notion of strength for the yield seeking motive relative to risk aversion, namely the ratio $\kappa = \alpha/\gamma \in [0, \infty)$.

Given that, we characterize the effect of a change in the interest rate as our main theoretical result in Proposition 2.

PROPOSITION 2. If the monetary authority decreases i_1:

1. Unconstrained firms do not change $\rho^*(A)$. Moreover, $\exists\ \varphi > 0$ such that if $\kappa > \varphi$, unconstrained firms increase their LT debt issuance, i.e.:

$$\frac{\partial d_l(A)}{\partial i_1} < 0 \quad \text{for } A \in (\overline{A}, I]$$

2. If $\kappa \rightarrow \infty$, then $\frac{\partial d_l(A)}{\partial i_1} \rightarrow -\infty$ for $A \in (\overline{A}, I]$; and $\frac{\partial \frac{\partial d_s(A)}{\partial i_1}}{\partial \kappa} = 0\ \forall A$.

3. Constrained firms increase $\rho^*(A)$ and reduce short-term debt. LT debt is unchanged.

Proof. See the Theory Appendix. □

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[33] A model with moral hazard frictions and risk-neutral investors, mimicking the original Holmström & Tirole (1998, 2000)'s framework, yields equally counterfactual predictions. Namely, in reaction to an expansionary interest rate shock by the FED, small firms lengthen debt maturity, whereas large ones do not adjust at all. This confirms the intuition that in models with simple credit frictions, constrained companies are in general more reactive to monetary policy, and the introduction of debt maturity just adds one credit margin along which this fact materializes (see discussion in section 1.1). A full derivation and description of this model is in the Theory Appendix.

The first condition on κ requires that yield-seeking motives are strong relative to risk-aversion. From the perspective of the model dynamics, this implies that demand functions for LT bonds react substantially to variations in the policy rate and that it is also relatively elastic. Under this condition, the upward demand shift due to a monetary interest rate loosening creates a mismatch with the existing supply of LT bonds. Constrained companies already issue LT debt at their limit, hence only large unconstrained firms can accommodate the demand shift.

The second result in Proposition 2 clarifies that, under κ sufficiently large, the effect on LT debt of unconstrained firms will dominate any adaptation of constrained firms in magnitude. In turn, this means that the model accommodates variations of the LT debt share of large companies which are arbitrarily larger than those of small companies, thereby matching the cross-sectional empirical evidence. Moreover, as both small and large companies adjust in the direction of lengthening the debt maturity, the model also aligns with aggregate-level empirical facts.

1.5 Empirical Evidence on the Model's Mechanism

Our model is consistent with the empirical results on the relation between the FED interest rate policy and corporate debt maturity. Relative to a basic framework with credit frictions only, the key ingredient for aligning the model with the data is the inclusion of yield-oriented investors. In particular, a mechanism arises whereby, in the aftermath of a rate cut, there is a boost in demand for LT bonds by such yield-seeking investors. Next, credit frictions imply that unconstrained (i.e., large) companies can accommodate such upward demand shifts, whereas smaller firms are at their debt-limit already and hence cannot issue LT bonds.

We test such mechanism. In practical terms, this requires showing that, in reaction to a monetary rate loosening, yield-seeking is associated to increased holdings of corporate LT bonds and that large companies increase issuance of such debtsecurities relatively more. However, in absence of further evidence, these adjustments may be driven by demand or supply motives. For dissecting their relative contribution, we additionally look at the price of the newly issued debt (i.e., the coupon rate). In this respect, to the extent that large companies also experience a relative stronger reduction in financing rates, it can be argued that their adjustments are relatively more demand-driven.[34]

[34] For bonds, an inverse relation holds between price and interest rate, that is, bond prices go up when the interest (coupon) rate goes down. Hence, a joint increase in issuance and reduction in interest rate describes a demand-driven adjustment, as it also corresponds to a boost in bond prices.

In the rest of this section, we first present the empirical analysis of corporate bonds mutual funds holdings and then report results on corporate debt issuance.

1.5.1 Monetary Policy and Corporate Bonds Mutual Funds Holdings

We employ once again local projections to analyze the dynamic response of CBMF to monetary interest rate variations:

$$\Delta y_{m,t+h} = \beta_{1,h}\Delta EFFRt + \beta_{2,h}HYm + \beta_{3,h}\Delta EFFRt * HY_m + \Gamma_h X_{m,t-1} + \mu_m + \mu_t + e_{m,t+h} \quad (15)$$

The dependent variable, $\Delta y_{m,t+h}$, is given by the growth between year-quarter $t-1$ and $t+h$ of fund m log volume of corporate bond holdings (or log portfolio's average weighted maturity).[35] Our coefficient of interest is $\beta_{3,h}$, loading the interaction between the quarterly EFFR variation, $\Delta EFFR_t$, and a dummy, HY_m, with value 1 if fund m is high-yield, our proxy for yield-seeking mutual funds. Importantly, $\beta_{3,h}$ captures the relative response of high-yield mutual funds' portfolios (as compared to investment grade ones) to variation in the interest rate policy of the FED, therefore sizing impact of reach-for-yield motives on such relation. $X_{m,t-1}$ is a vector of time-varying fund-level controls, including the interaction of: i) HY_m with several macro-level controls; ii) other lagged fund characteristics (turnover ratio, expense ratio, log asset size and returns) with $\Delta EFFR_t$. Moreover, we augment the model with fund and year-quarter fixed effects (μ_m and μ_{yq}), respectively controlling for time-invariant heterogeneity at the level of the fund and for common shocks across all funds in a given year-quarter. $e_{m,t+h}$ is an error term, double-clustered at the fund and year-quarter level.

To start with, Figure 11 describes the relative response of HY-funds' corporate bonds holdings to a 25 b.p. cut to the EFFR. The effect is markedly positive at impact and peaks up 6 quarters after the shock. In particular, at impact, HY-funds increase their corporate bonds holdings by 2.87 p.p. (and by 6 p.p. one year after the shock). We find a similar relative jump at impact when using exogenous monetary policy shocks, though the effect appears much less persistent and reverts back to zero 2 quarters after the shock (Figure 12).

Next, Figure 13 confirms that, following an interest rate descent, and on top of buying corporate bonds, HY-funds additionally tilt their portfolio towards debtsecurities with longer maturity. In

[35] Ideally, we would like to look at the average weighted maturity of the fund's holdings of corporate bonds, rather than of the overall portfolio. Unfortunately, though, only the latter is available in the CRSP Survivor Bias Free dataset.

detail, one year after the 25 b.p. reduction in the policy rate, portfolio maturity goes up by 2 p.p.. A similar effect emerges in Figure 14, where we exploit exogenous monetary policy shocks rather than the raw EFFR quarterly variations.

1.5.2 Monetary Poilicy and Corporate Bonds Issuance

EXTENSIVE MARGIN: FREQUENCY OF ISSUANCE

First, we check whether the likelihood of issuing new LT bonds is differently affected by interest rate changes across small and large companies. To this end, we resort to Model 2, i.e., a panel model saturated with firm and industry-time fixed effects and in which the interaction between the large-firm dummy and the interest rate change is horse-raced against different balance-sheet channels. The dependent variable is $1(\text{Issue})_{f,t+h}$, a dummy variable with value 1 if firm f issues LT bonds in yearquarter t + h and with value 0 if it does not. Hence, at horizon h, the coefficient $\beta_{3,h}$ measures the relative difference in large companies' probability of issuing new debt as of year-quarter t + h, induced by a FED revision of the policy rate at t.

Results are displayed in Figure 15. The comparative response of large companies (vis-a-vis smaller ones) to a 25 b.p. cut in the EFFR is markedly more positive at impact, when it amounts to 28 b.p.. Such increase corresponds to an additional 4% jump relative to the average likelihood of issuing LT bonds as of time t. The effect extends over time, peaking 2 year-quarters after the rate change and vanishing in roughly one year and a half. To understand the absolute impact of monetary policy on the likelihood of issuing LT bonds, we report in Table 3 regressions with the dependent measured as of year-quarter t and t + 1 and which do not include timevarying fixed effects (columns 1 and 4, respectively). Indeed, this exercise suggests that smaller companies also increase the likelihood of issuing bonds when the policy rate falls, however with a magnitude which is twice as small as that observed for larger firms.

We perform different robustness checks. In Figure 16 we replicate the analysis using the familiar exogenous monetary policy shocks by Gürkaynak et al. (2005) and Jarocinski & Karadi (2020) and obtain similar results. As we exploit interest rate shocks - more prevalent before the last Global Financial Crisis - we additionally test whether our findings survive the exclusion from our sample of the observations from 2009 onward. Appendix Figures A8 and A9 suggest that the response is qualitatively comparable and, if anything, quantitatively larger.

INTENSIVE MARGIN: FINANCING COSTS

We aim to understand the cross-sectional differences in the reaction of financing costs to variations in the FED's interest rate policy. One issue with this analysis is that only a tiny subset of companies ever issue bonds in two consecutive quarters. In operative terms, this means that if we were to apply a first-differenced model, we would be left with very few observations and, ultimately, a meaningless crosssectional comparisons across a very small set of companies (highly skewed towards large firms).

Hence, we rather resort to the following model in levels:

$$y_{f,t+h} = \beta_{1,h} EFFR_t + \beta_{2,h} Large_{f,t-1} + \beta_{3,h} Large_{f,t-1} * EFFR_t + \Phi_h X_{f,t-1} + \mu_f + \mu_{s,t} + \xi_{f,t+h} \quad (16)$$

As dependent variable, we use $Coupon_{f,t+h}$, i.e., the coupon rate on firm f's newly issued LT bonds at year-quarter t + h. This is regressed against the current level of the policy rate, $EFFR_t$, interacted with the dummy for large firms, $Large_{f,t-1}$. The vector of controls $X_{f,t-1}$ aligns to the previously used empirical models.[36] As usual, we augment the model with firm and industry*year-quarter fixed effects (μ_f and $\mu_{s,t+h}$) and double-cluster the error term, $\xi_{f,t+h}$, accordingly.

Figure 17 reports the estimated coefficients $\beta_{3,h}$, calibrated to a 1 p.p. lower EFFRlevel. Clearly, when the EFFR is lower, large firms' coupon rate is also lower (as compared to that of smaller companies). Quantitatively speaking, a 1 p.p. looser EFFR grants big companies an additional reduction in the coupon rate by roughly 10 b.p.., resulting in a further 1.6% cut in financing costs relatively to their average level. Table 4 describes the background regressions for h = 0, 1, also shown with different set of fixed effects allowing to evaluate the absolute response of the coupon rate. In columns 1 and 3, in fact, we report a regression which, differently from the model in Equation (16), excludes time-varying fixed effects. Both in year-quarter t and t + 1, the relation between the coupon rate and the EFFR is generally positive across smaller firms, too. Put differently, based on this exercise, for all companies the coupon rate goes down when the EFFR is lower, but the effect is stronger among large firms.

As clear from Table 4, there is a significant loss of observations due to the application of (3-digit SIC)-industry*year-quarter fixed effects. Bond issuance at the firm-level is indeed quite lumpy

[36] That is, $X_{f,t-1}$ includes the full interaction of: $Large_{f,t-1}$ and macro controls; firm-level controls and $EFFR_t$. Relatively to previous models, we report few macro-controls (term-spread, corporate spread and share of Treasuries with maturity above 20-year) in levels rather than in first-differences.

across time. Therefore, narrowing the comparison within granular industries implies the loss of many (within-industry) singletons. One additional problem is that in this framework the within-industry cross-sectional comparison will comprehend few firms. Hence, for robustness purposes, we apply looser industry definitions and check that our findings go through. In Table 5, in columns 1-4, we replicate our analysis, but this time comparing all companies with each other. In this case, the $Large_{f,t-1}$ dummy captures those companies in the topquartile of the entire sample of NFCs in the US stock market. The result that large companies' coupon rates descent along with the EFFR still goes through. A similar pattern emerges when applying increasingly more granular industry-definition (sectoral-level in columns 5-8 and 2-digit SIC industry in columns 9-12).

1.6 Conclusions

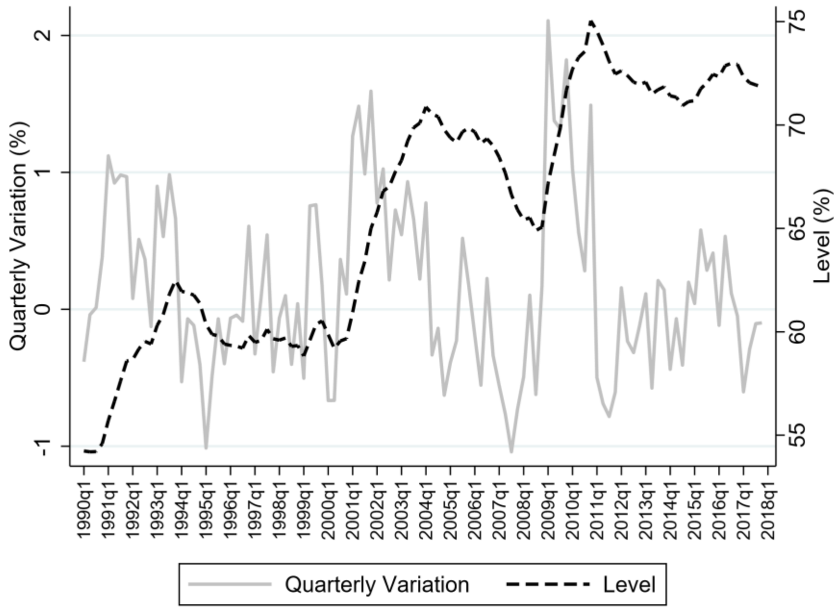
Firms' resilience to shocks crucially depends on the maturity structure of corporate debt. Hence, understanding whether and how monetary policy affects firms' debt maturity is key for gauging the implications for firms' risk of central banks' policies.

This paper provides novel empirical evidence on the relation between the interest rate policy by the FED and the maturity structure of the US corporate (non-financial) sector. Our robust findings suggest that, following a policy rate cut, firms lengthen debt maturity. The effect is entirely driven by very large firms, whereas smaller companies' debt maturity is generally not responsive to monetary policy. Existing theoretical frameworks do not provide an adequate explanation for these findings. Hence, we build a model combining credit frictions due to firms' moral hazard and yield-oriented investors - who increase the demand for long-term bonds when the interest rate goes down. Only large and unconstrained companies can accommodate such upward shift in demand, so that the model aligns with the empirical evidence. We bring the model mechanism to the data and find supportive evidence. As a matter of fact, following a policy rate decline: i) relatively more yield-oriented mutual funds increase their holdings of corporate bonds and tilt their portfolio towards longer-term debt securities; ii) large firms issue more debt and at lower rates, indicating that their adjustment is demand driven.

Ultimately, our work highlights how monetary policy impacts the maturity profile of very large corporations, whose dynamics have significant consequences for the business cycle (Crouzet & Mehrotra forthcoming). An important open question - left for future research - is whether the documented interaction between monetary policy and corporate debt maturity has implications for business cycle and systemic risk.

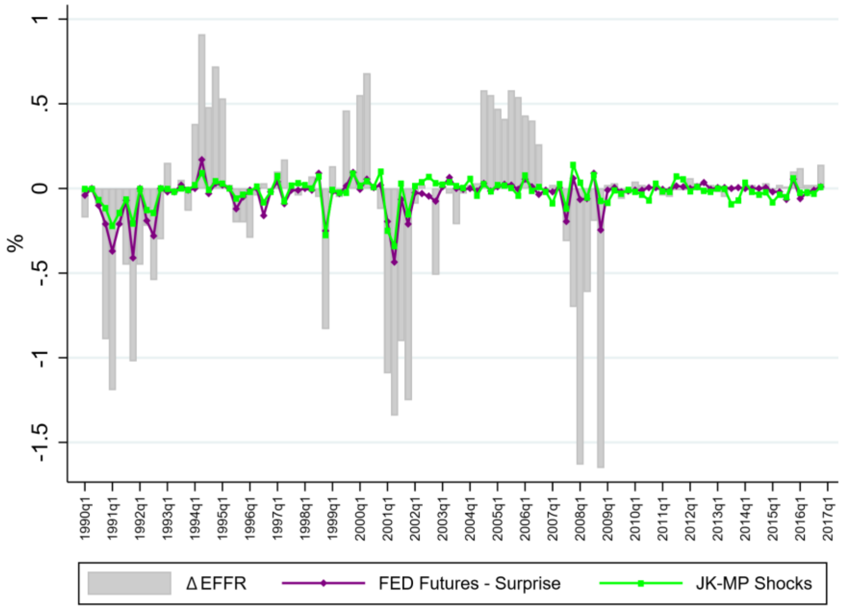
Figures

FIGURE 1: % of LT-Dedt – Aggregate Level



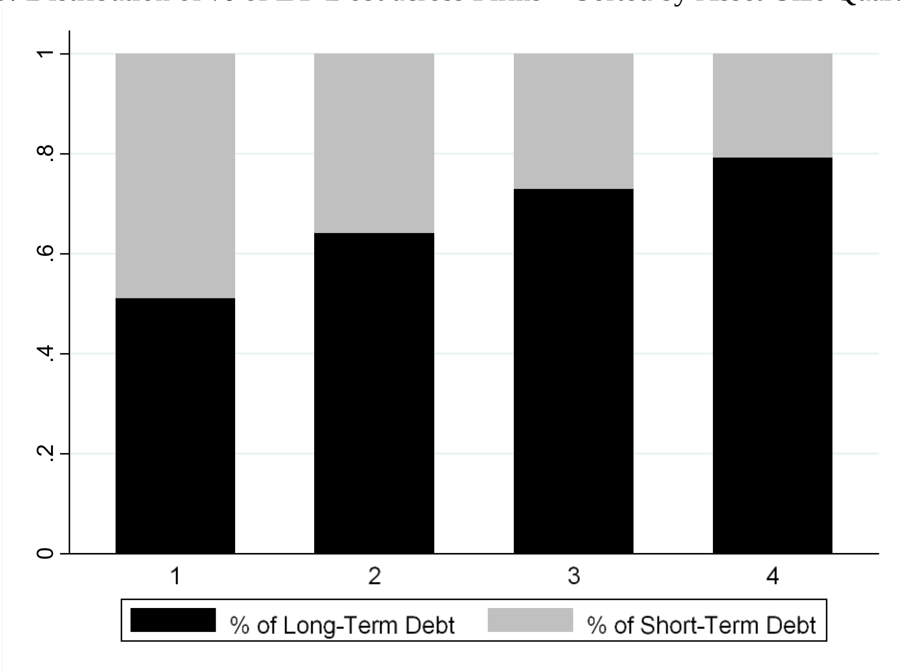
This figure shows the evolution of the aggregate share of LT debt (i.e. with outstanding maturity above 1 year). The black dashed line reports the series in levels, the grey one in first-differences. Following Greenwood et al. (2010), LT debt is defined as the sum of corporate bonds and mortgages and industrial revenues. The remaining short-term corporate debt is proxied by the sum of shortterm loans (and advances) and commercial paper.

FIGURE 2: Measures of Changes in the Policy Rate



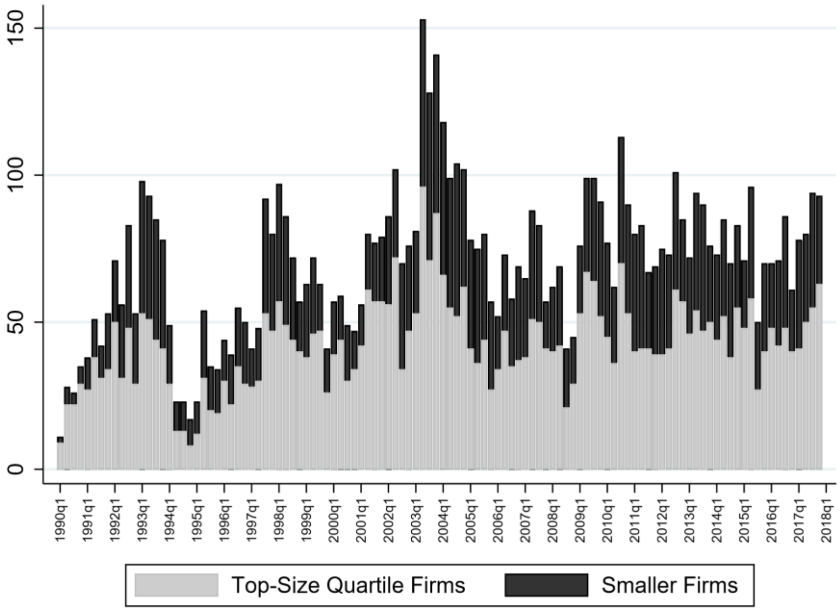
The grey bars show the quarterly variation of the Effective FED Funds Rate. The purple (green) solid line, connected by diamonds (squares), reports the Gürkaynak et al. 2005 (Jarocinski & Karadi 2020) monetary policy shocks.

FIGURE 3: Distribution of % of LT-Debt across Firms – Sorted by Asset-Size Quartiles



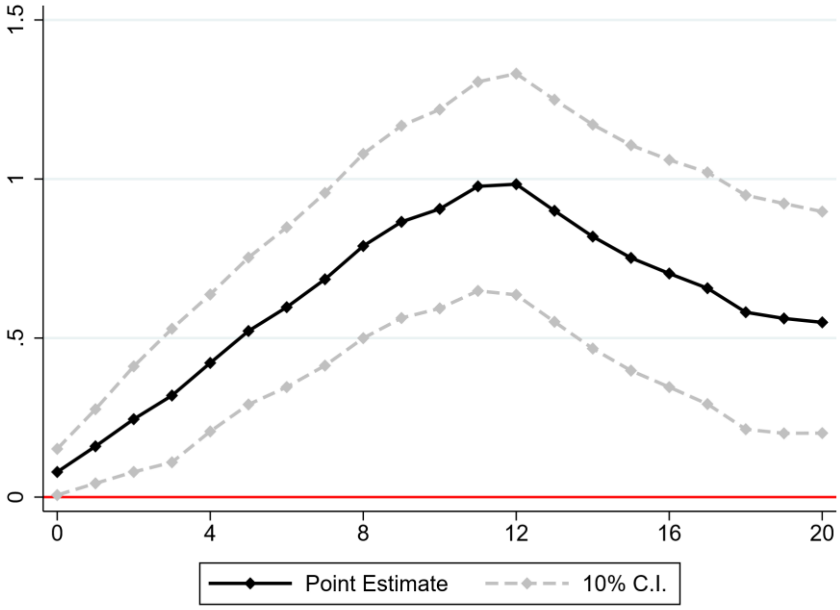
This chart shows the average % of LT-Debt across different groups of companies (black bars); the complement to 100% gives the average % of ST-Debt (grey bars). Firms are sorted according to quartiles of their 3-digit SIC industry asset-size distribution. Quartiles are reported on the x-axis.

FIGURE 4: Bond Issuance Over Time



This chart shows the number of bond issuances over time. Companies are sorted based on their (3digit SIC) asset-size distribution. The grey bars report the number of bond issuances by companies in the upper quartile; the black bars by all other companies. The vertical sum of the grey and black bars provides the total number of bond issuances per year-quarter.

FIGURE 5: Monetary Policy and Debt Maturity Structure: Aggregate Response

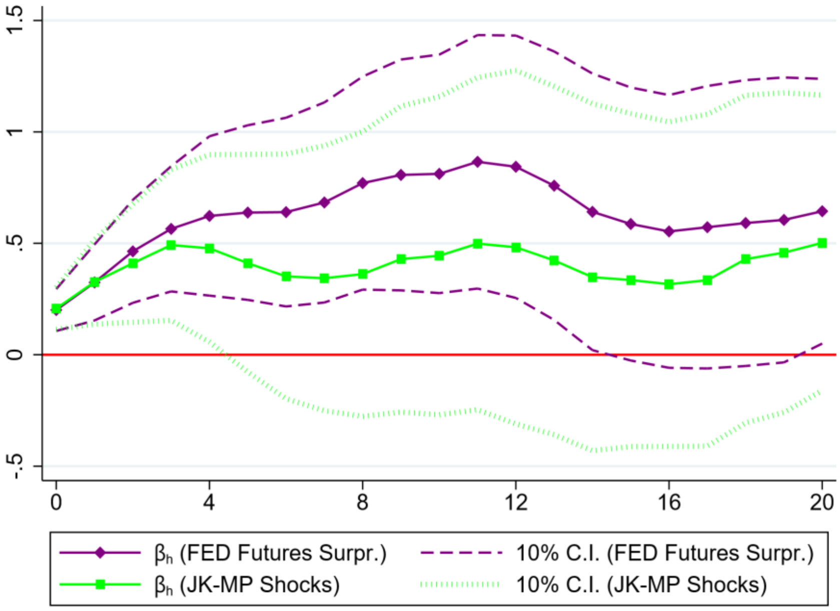


This figure depicts the response of the aggregate-level share of LT debt to a 25 b.p. cut in the EFFR. Formally, it shows the coefficients $\beta_{1,h}$ from the estimation of the following local projection model:

$$\Delta_h y_{t+h} = \beta_{1,h} \Delta EFFR_t + MacroControls_{t-1} + u_{t,h}$$

The dependent variable, $\Delta_h y_{t+h}$, represents the growth of the LT-debt share (expressed in p.p.) from year-quarter $t-1$ to year-quarter $t+h$. $\Delta EFFR_t$ is the quarterly EFFR change. $MacroControls_{t-1}$ is a vector of lagged macroeconomic controls, including annual GDP growth and inflation rate, a dummy for recessions, the quarterly variation in the share of LT treasuries, in the corporate spread and in the 10y-3m term-spread. $u_{t,h}$ is a robust error-term. The x-axis is measured in terms of quarters after the shock. The black solid line reports the point estimates for $\beta_{1,h}$; the dashed grey line the 10% confidence intervals.

FIGURE 6: Monetary Policy and Debt Maturity Structure: Aggregate Response using Exogenous Shocks

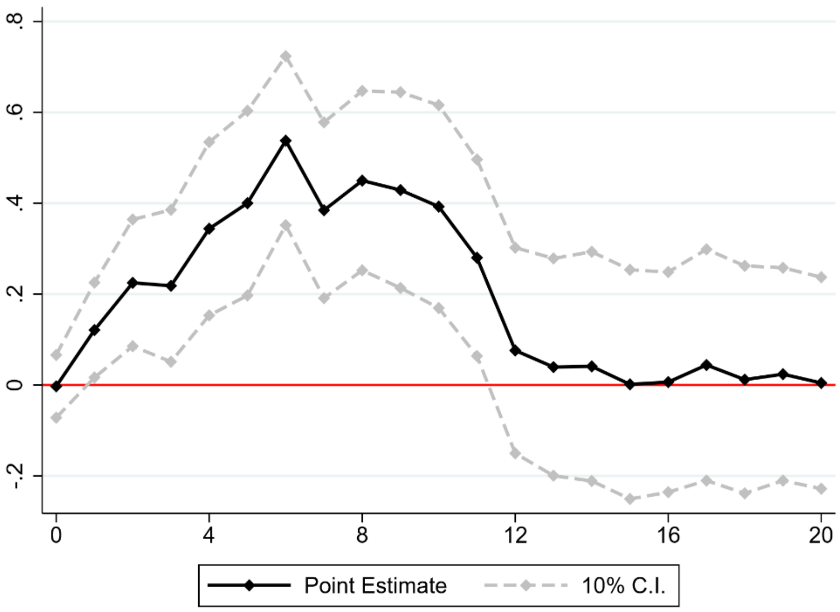


This figure depicts the response of the aggregate-level share of LT debt to a 1 s.d. reduction in the monetary policy shock. Formally, it shows the coefficients $\beta_{1,h}$ from the estimation of the following local projection model:

$$\Delta_h y_{t+}h = \beta_{1,h}\varepsilon^{mp}{}_t + MacroControls_{t-1} + u_{t,h}$$

The dependent variable, $\Delta_h y_{t+h}$, represents the growth of the LT-debt share (expressed in p.p.) from year-quarter $t - 1$ to year-quarter $t + h$. $\varepsilon^{mp}{}_t$ is an exogenous monetary policy shock, gathered from either Gürkaynak et al. (2005) or from Jarocinski & Karadi (2020). $MacroControls_{t-1}$ is a vector of lagged macroeconomic controls, including annual GDP growth and inflation rate, a dummy for recessions, the quarterly variation in the share of LT treasuries, in the corporate spread and in the 10y-3m term-spread. $u_{t,h}$ is a robust error-term. The x-axis is measured in terms of quarters after the shock. The purple (green) solid line, connected by diamonds (squares), reports the point estimates for $\beta_{1,h}$ using the Gürkaynak et al. 2005 (Jarocinski & Karadi 2020) shocks and the dashed (dotted) purple (green) line the respective 10% confidence intervals.

FIGURE 7: Monetary Policy and Debt Maturity Structure – Relative Response of Large Firms

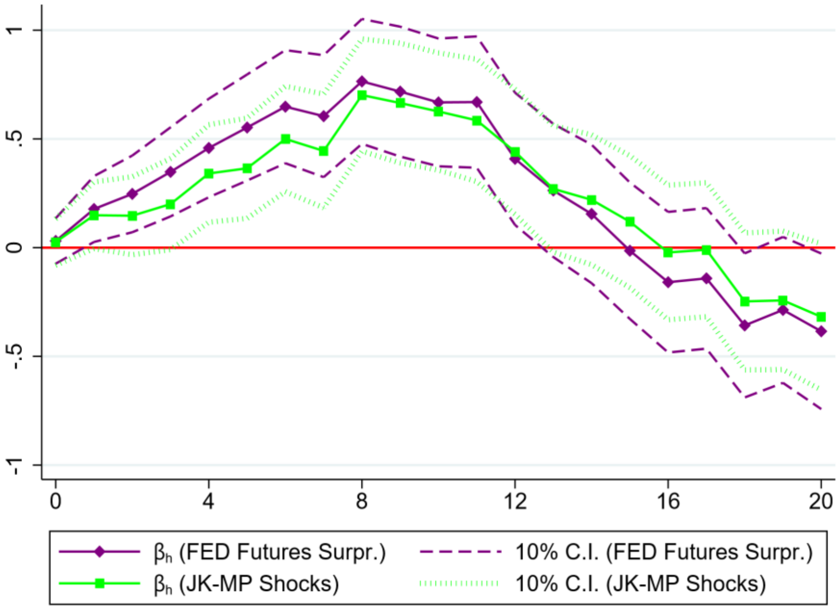


This figure depicts the relative response of companies in the top quartile of the 3-digit SIC industry asset-size distribution to a 25 b.p. cut in the EFFR (as compared to smaller firms). Formally, it shows the coefficients $\beta_{3,h}$ from the estimation of the following local projection model:

$$\Delta_h y_{f,t+h} = \beta_{1,h}\Delta EFFR_t + \beta_{2,h} Large_{f,t-1} + \beta_{3,h} Large_{f,t-1} * \Delta EFFR_t + X_{f,t-1} + \mu_f + \mu_{s,t} + u_{t,h}$$

The dependent variable, $\Delta_h y_{f,t+h}$, represents the growth of the share of LT-debt - expressed in p.p. from year-quarter $t - 1$ to year-quarter $t + h$. $\Delta EFFR_t$ is the quarterly change in the EFFR. $Large_{f,t-1}$ is a dummy variable with value 1 a firm is in the top-quartile of the industry-wide asset-size distribution, and 0 otherwise. $X_{f,t-1}$ is a vector of controls, including the interaction of $Large_{f,t-1}$ with several macro-controls and of $\Delta EFFR_t$ with other firm characteristics, namely lagged sales growth, leverage and liquid assets. μ_f and $\mu_{s,t}$ represent vectors of firm and industry*year-quarter fixed effects, respectively. $u_{t,h}$ is an error-term, double-clustered at the firm and industry*year-quarter level. The x-axis is measured in terms of quarters after the shock. The black solid line reports the point estimates for $\beta_{3,h}$; the dashed grey line the 10% confidence intervals.

FIGURE 8: Monetary Policy and Debt Maturity Structure – Relative Response of Large Firms using Exogenous Shocks

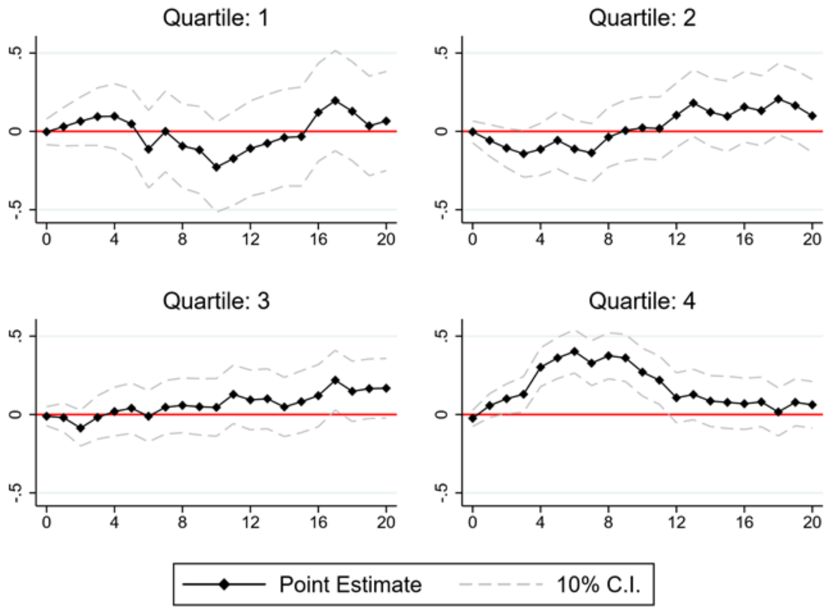


This figure depicts the relative response of companies in the top quartile of the 3-digit SIC industry asset-size distribution to a 1 s.d. b.p.reduction in monetary policy shock (as compared to smaller firms). Formally, it shows the coefficients $\beta_{3,h}$ from the estimation of the following local projection model:

$$\Delta_h y_{f,t+h} = \beta_{1,h}\varepsilon^{mp}_t + \beta_{2,h}Large_{f,t-1} + \beta_{3,h}Large_{f,t-1} * \varepsilon^{mp}_t + X_{f,t-1} + \mu_f + \mu_{s,t} + u_{t,h}$$

The dependent variable, $\Delta_h y_{f,t+h}$, represents the growth of the share of LT-debt - expressed in p.p. from year-quarter $t-1$ to year-quarter $t+h$. ε^{mp}_t is an exogenous monetary policy shock, derived either from Gürkaynak et al. (2005) or from Jarocinski & Karadi (2020). $Large_{f,t-1}$ is a dummy variable with value 1 a firm is in the top-quartile of the industry-wide asset-size distribution, and 0 otherwise. $X_{f,t-1}$ is a vector of controls, including the interaction of $Large_{f,t-1}$ with several macro-controls and of . ε^{mp}_t with other firm characteristics, namely lagged sales growth, leverage and liquid assets. μ_f and $\mu_{s,t}$ represent vectors of firm and industry*year-quarter fixed effects, respectively. $u_{t,h}$ is an errorterm, double-clustered at the firm and industry*year-quarter level. The x-axis is measured in terms of quarters after the shock. The purple (green) solid line, connected by diamonds (squares), reports the point estimates for $\beta_{1,h}$ using the Gürkaynak et al. 2005 (Jarocinski & Karadi 2020) shocks and the dashed (dotted) purple (green) line the respective 10% confidence intervals.

FIGURE 9: Monetary Policy and Debt Maturity Structure – Absolute Firm-Level Response

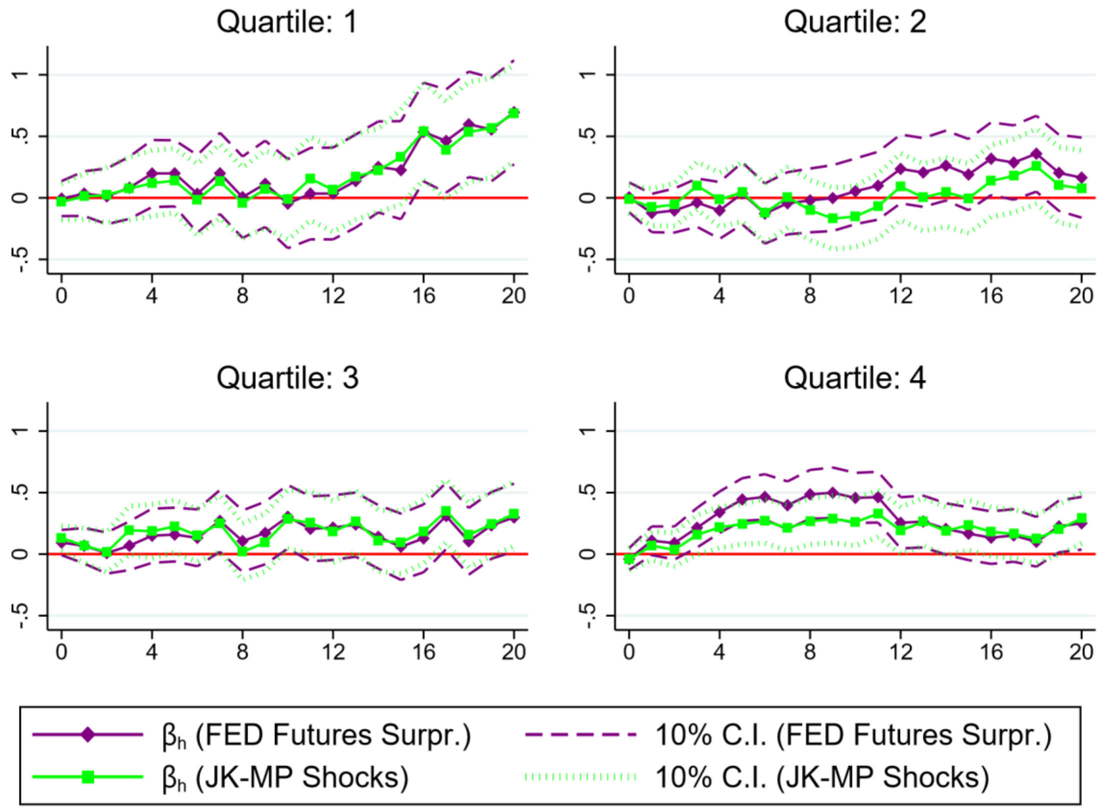


This figure depicts the absolute response of companies in different size quartiles of the 3-digit SIC industry asset-size distribution to a 25 b.p. cut in the EFFR. Formally, it shows the coefficients $\beta_{1,h}$ from the estimation of the following local projection model:

$$\Delta_h y_{f,t+h} = \beta_{1,h}\Delta EFFR_t + MacroControls_{t-1} + X_{f,t-1} + \mu_f + \mu_{s,t} + u_{t,h}$$

The dependent variable, $\Delta_h y_{f,t+h}$, represents the growth of the LT-debt share (expressed in p.p.) from year-quarter t − 1 to year-quarter t + h. $\Delta EFFR_t$ is the quarterly EFFR change. $MacroControls_{t-1}$ is a vector of lagged macroeconomic controls, including annual GDP growth and inflation rate, a dummy for recessions, the quarterly variation in the share of LT treasuries, in the corporate spread and in the 10y-3m term-spread. $X_{f,t-1}$ is a vector of firm-level controls, including lagged sales growth, leverage and liquid assets. μ_f is a vector of firm fixed effects. $u_{t,h}$ is an error-term, double-clustered at the firm and industry*year-quarter level. The x-axis is measured in terms of quarters after the shock. The black solid line reports the point estimates for $\beta_{1,h}$; the dashed grey line the 10% confidence intervals.

FIGURE 10: Monetary Policy and Debt Maturity Structure – Absolute Firm-Level Response using Exogenous Shocks

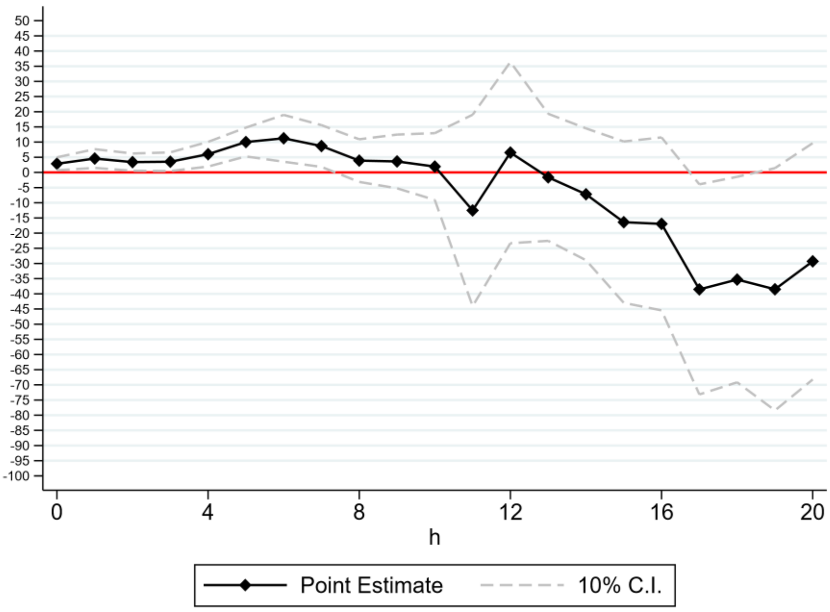


This figure depicts the absolute response of companies in different size quartiles of the 3-digit SIC industry asset-size distribution to a 1 s.d. reduction in monetary policy shock. Formally, it shows the coefficients $\beta_{1,h}$ from the estimation of the following local projection model:

$$\Delta_h y_{f,t+h} = \beta_{1,h}\varepsilon^{mp}_t + \text{MacroControls}_{t-1} + X_{f,t-1} + \mu_f + \mu_{s,t} + u_{t,h}$$

The dependent variable, $\Delta_h y_{f,t+h}$, represents the growth of the LT-debt share (expressed in p.p.) from year-quarter $t-1$ to year-quarter $t+h$. ε^{mp}_t is an exogenous monetary policy shock, gathered from either Gürkaynak et al. (2005) or from Jarocinski & Karadi (2020). $\text{MacroControls}_{t-1}$ is a vector of lagged macroeconomic controls, including annual GDP growth and inflation rate, a dummy for recessions, the quarterly variation in the share of LT treasuries, in the corporate spread and in the 10y3m term-spread. $X_{f,t-1}$ is a vector of firm-level controls, including lagged sales growth, leverage and liquid assets. μ_f is a vector of firm fixed effects. $u_{t,h}$ is an error-term, double-clustered at the firm and industry*year-quarter level. The x-axis is measured in terms of quarters after the shock. The purple (green) solid line, connected by diamonds (squares), reports the point estimates for $\beta_{1,h}$ using the Gürkaynak et al. 2005 (Jarocinski & Karadi 2020) shocks and the dashed (dotted) purple (green) line the respective 10% confidence intervals.

FIGURE 11: Monetary Policy and CBMFS' Corporate Bonds Holdings

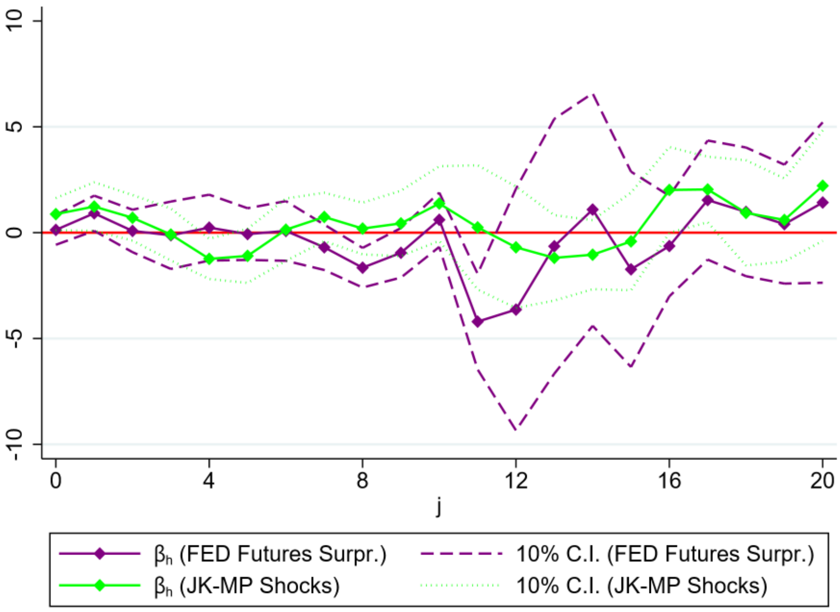


This figure shows the relative response of High-Yield (HY) corporate bonds mutual funds to a 25 b.p. cut in the EFFR (as compared to Investment-Grade funds). Formally, it shows the coefficients $\beta_{3,h}$ from the estimation of the following local projection model:

$$\Delta y_{m,t+h} = \beta_{1,h}\Delta EFFR_t + \beta_{2,h}HY_m + \beta_{3,h}\Delta EFFR_t * HY_m + \Gamma_h X_{m,t-1} + \mu_m + \mu_t + e_{m,t+h}$$

The dependent variable, $\Delta y_{m,t+h}$, is given by the growth between year-quarter $t-1$ and $t+h$ of fund m log volume of corporate bond holdings, expressed in p.p.. $\Delta EFFR_t$ gives the quarterly EFFR variation. HY_m is a dummy with value 1 if fund m is high-yield and with value 0 otherwise. $X_{m,t-1}$ is a vector of time-varying fund-level controls. It includes the full interaction of HY_m with a vector of macro-level controls, namely annual GDP growth and inflation rate, a dummy for recessions, the quarterly variation in the share of LT treasuries, in the corporate spread and in the 10y-3m termspread. Moreover, we controls for the full interaction of other lagged fund characteristics (turnover ratio, expense ratio, log asset size and returns) with $\Delta EFFR_t$. μ_m and μ_{yq} are fund- and year-quarter fixed effects. $e_{m,t+h}$ is an error term, double-clustered at the fund and year-quarter level. The black solid line reports the point estimates for $\beta_{3,h}$; the dashed grey line the 10% confidence intervals.

FIGURE 12: Monetary Policy and CBMFS' Corporate Bond Holdings using Exogenous Shocks

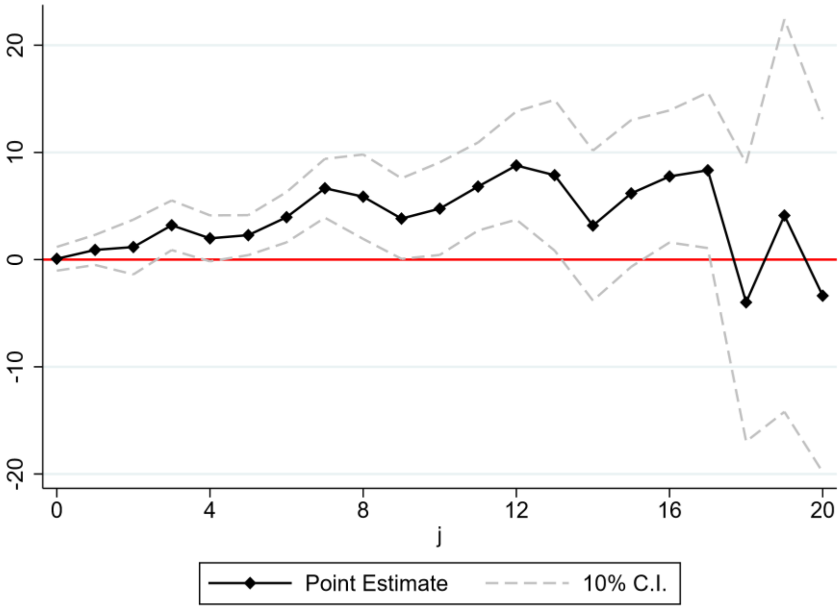


This figure shows the relative response of High-Yield (HY) corporate bonds mutual funds to a 25 b.p. cut in the EFFR (as compared to Investment-Grade funds). Formally, it shows the coefficients $\beta_{3,h}$ from the estimation of the following local projection model:

$$\Delta y_{m,t+h} = \beta_{1,h}\,\varepsilon^{mp}_t + \beta_{2,h} HY_m + \beta_{3,h}\varepsilon^{mp}_t * HY_m + \Gamma_h X_{m,t-1} + \mu_m + \mu_t + e_{m,t+h}$$

The dependent variable, $\Delta y_{m,t+h}$, is given by the growth between year-quarter $t-1$ and $t+h$ of mp fund m log corporate bond holdings, expressed in p.p.. ε_t is an exogenous monetary policy shock, gathered from either Gürkaynak et al. (2005) or from Jarocinski & Karadi (2020). HY_m is a dummy with value 1 if fund m is high-yield and with value 0 otherwise. $X_{m,t-1}$ is a vector of time-varying fund-level controls. It includes the full interaction of HY_m with a vector of macro-level controls, namely annual GDP growth and inflation rate, a dummy for recessions, the quarterly variation in the share of LT treasuries, in the corporate spread and in the 10y-3m term-spread. Moreover, we controls for the full interaction of other lagged fund characteristics (turnover ratio, expense ratio, log mp asset size and returns) with ε^{mp}_t. μ_m and μ_{yq} are fund- and year-quarter fixed effects. $e_{m,t+h}$ is an error term, double-clustered at the fund and year-quarter level. The purple (green) solid line, connected by diamonds (squares), reports the point estimates for $\beta_{1,h}$ using the Gürkaynak et al. 2005 (Jarocinski & Karadi 2020) shocks and the dashed (dotted) purple (green) line the respective 10% confidence intervals.

FIGURE 13: Monetary Policy and CBMFS' Portfolio Maturity

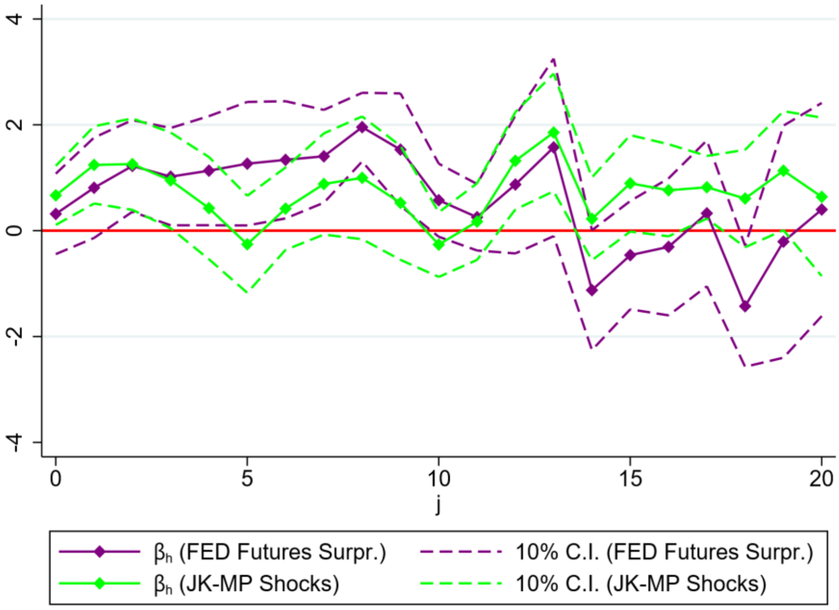


This figure shows the relative response of High-Yield (HY) corporate bonds mutual funds to a 25 b.p. cut in the EFFR (as compared to Investment-Grade funds). Formally, it shows the coefficients $\beta_{3,h}$ from the estimation of the following local projection model:

$$\Delta y_{m,t+h} = \beta_{1,h}\Delta EFFR_t + \beta_{2,h}HY_m + \beta_{3,h}\Delta EFFR_t * HY_m + \Gamma_h X_{m,t-1} + \mu_m + \mu_t + e_{m,t+h}$$

The dependent variable, $\Delta y_{m,t+h}$, is given by the growth between year-quarter $t-1$ and $t+h$ of fund m log (weighted) average maturity, expressed in p.p.. $\Delta EFFR_t$ gives the quarterly EFFR variation. HY_m is a dummy with value 1 if fund m is high-yield and with value 0 otherwise. $X_{m,t-1}$ is a vector of time-varying fund-level controls. It includes the full interaction of HY_m with a vector of macro-level controls, namely annual GDP growth and inflation rate, a dummy for recessions, the quarterly variation in the share of LT treasuries, in the corporate spread and in the 10y-3m term-spread. Moreover, we controls for the full interaction of other lagged fund characteristics (turnover ratio, expense ratio, log asset size and returns) with $\Delta EFFR_t$. μ_m and μ_{yq} are fund- and year-quarter fixed effects. $e_{m,t+h}$ is an error term, double-clustered at the fund and year-quarter level. The black solid line reports the point estimates for $\beta_{3,h}$; the dashed grey line the 10% confidence intervals.

FIGURE 14: Monetary Policy and CBMFS' Portfolio Maturity – using Exogenous Shocks

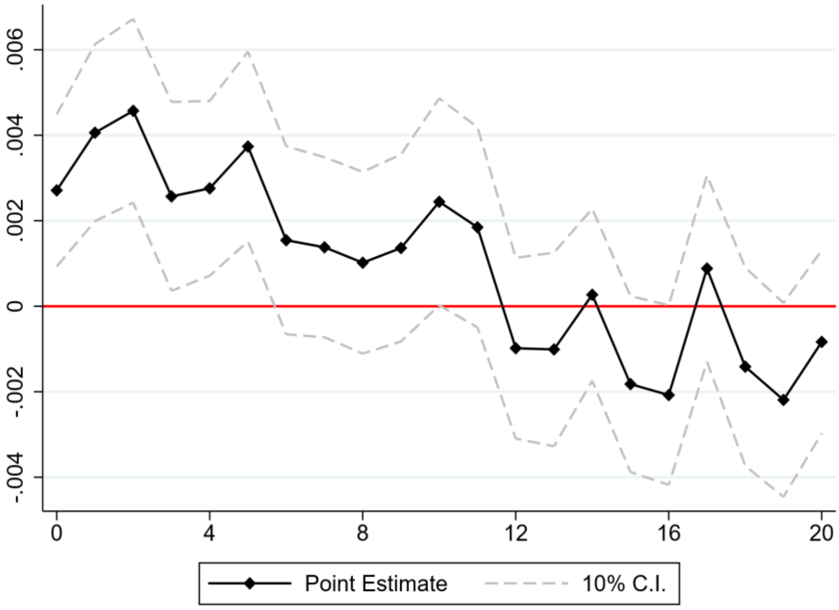


This figure shows the relative response of High-Yield (HY) corporate bonds mutual funds to a 25 b.p. cut in the EFFR (as compared to Investment-Grade funds). Formally, it shows the coefficients $\beta_{3,h}$ from the estimation of the following local projection model:

$$\Delta y_{m,t+h} = \beta_{1,h}\varepsilon^{mp}_t + \beta_{2,h}HY_m + \beta_{1,h}\varepsilon^{mp}_t * HY_m + \Gamma_h X_{m,t-1} + \mu_m + \mu_t + e_{m,t+h}$$

The dependent variable, $\Delta y_{m,t+h}$, is given by the growth between year-quarter t− 1 and t+ h of fund mp m log (weighted) average maturity, expressed in p.p.. ε_t is an exogenous monetary policy shock, gathered from either Gürkaynak et al. (2005) or from Jarocinski & Karadi (2020). HY_m is a dummy with value 1 if fund m is high-yield and with value 0 otherwise. $X_{m,t-1}$ is a vector of time-varying fund-level controls. It includes the full interaction of HY_m with a vector of macro-level controls, namely annual GDP growth and inflation rate, a dummy for recessions, the quarterly variation in the share of LT treasuries, in the corporate spread and in the 10y-3m term-spread. Moreover, we controls for the full interaction of other lagged fund characteristics (turnover ratio, expense ratio, log mp asset size and returns) with ε_t . μ_m and μ_{yq} are fund- and year-quarter fixed effects. $e_{m,t+h}$ is an error term, double-clustered at the fund and year-quarter level. The purple (green) solid line, connected by diamonds (squares), reports the point estimates for $\beta_{1,h}$ using the Gürkaynak et al. 2005 (Jarocinski & Karadi 2020) shocks and the dashed (dotted) purple (green) line the respective 10% confidence intervals.

FIGURE 15: Monetary Policy and Likelihood of Issuing LT Bonds – Relative Response of Large Companies

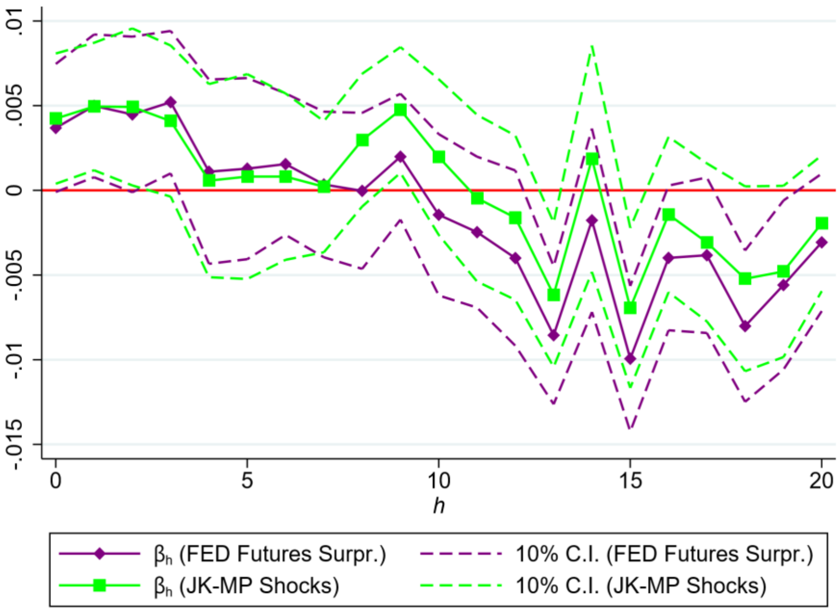


This figure depicts the relative response of companies in the top quartile of the 3-digit SIC industry asset-size distribution to a 25 b.p. cut in the EFFR (as compared to smaller firms). Formally, it shows the coefficients $\beta_{3,h}$ from the estimation of the following local projection model:

$$y_{f,t+h} = \beta_{1,h}\Delta EFFR_t + \beta_{2,h}Large_{f,t-1} + \beta_{3,h}\Delta EFFR_t * Large_{f,t-1} + X_{f,t-1} + \mu_f + \mu_{s,t} + u_{t,h}$$

The dependent variable, $y_{f,t+h}$, is a dummy with value 1 if firms f issues LT bonds in year-quarter t + h and with value 0 if it does not. $\Delta EFFR_t$ is the quarterly change in the EFFR. $Large_{f,t-1}$ is a dummy variable with value 1 a firm is in the top-quartile of the industry-wide asset-size distribution, and 0 otherwise. $X_{f,t-1}$ is a vector of controls, including the interaction of $Large_{f,t-1}$ with several macro-controls and of $\Delta EFFR_t$ with other firm characteristics, namely lagged sales growth, leverage and liquid assets. μ_f and $\mu_{s,t}$ represent vectors of firm and industry*year-quarter fixed effects, respectively. $u_{t,h}$ is an error-term, double-clustered at the firm and industry*year-quarter level. The x-axis is measured in terms of quarters after the shock. The black solid line reports the point estimates for $\beta_{3,h}$; the dashed grey line the 10% confidence intervals.

FIGURE 16: Monetary Policy and Likelihood of Issuing LT Bonds – Relative Response of Large Companies using Exogenous Shocks

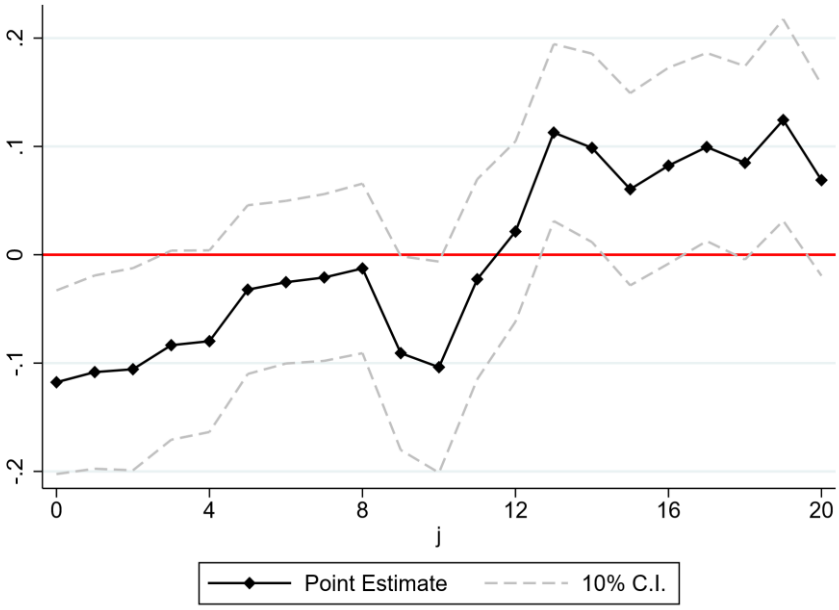


This figure depicts the relative response of companies in the top quartile of the 3-digit SIC industry asset-size distribution to a 25 b.p. cut in the EFFR (as compared to smaller firms). Formally, it shows the coefficients $\beta_{3,h}$ from the estimation of the following local projection model:

$$y_{f,t+h} = \beta_{1,h}\varepsilon^{mp}_t + \beta_{2,h}Large_{f,t-1} + \beta_{3,h}Large_{f,t-1} * \varepsilon^{mp}_t + X_{f,t-1} + \mu_f + \mu_{s,t} + u_{t,h}$$

The dependent variable, $y_{f,t+h}$, is a dummy with value 1 if firms f issues LT bonds in year-quarter and with value 0 if it does not. ε^{mp}_t is an exogenous monetary policy shock, gathered from either Gürkaynak et al. (2005) or from Jarocinski & Karadi (2020). $MacroControls_{t-1}$ is a vector of lagged macroeconomic controls, including annual GDP growth and inflation rate, a dummy for recessions, the quarterly variation in the share of LT treasuries, in the corporate spread and in the 10y-3m termspread. $X_{f,t-1}$ is a vector of firm-level controls, including lagged sales growth, leverage and liquid assets. μ_f is a vector of firm fixed effects. $u_{t,h}$ is an error-term, double-clustered at the firm and industry*year-quarter level. The x-axis is measured in terms of quarters after the shock. The purple (green) solid line, connected by diamonds (squares), reports the point estimates for $\beta_{1,h}$ using the Gürkaynak et al. 2005 (Jarocinski & Karadi 2020) shocks and the dashed (dotted) purple (green) line the respective 10% confidence intervals.

FIGURE 17: Monetary Policy and Financing Costs of LT Bonds – Relative Response of Large Companies



This figure depicts the relative response of companies in the top quartile of the 3-digit SIC industry asset-size distribution to a 1 p.p. lower EFFR (as compared to smaller firms). Formally, it shows the coefficients $\beta_{3,h}$ from the estimation of the following local projection model:

$$y_{f,t+h} = \beta_{1,h}EFFR_t + \beta_{2,h}Large_{f,t-1} + \beta_{3,h}Large_{f,t-1} * EFFR_t + \Phi X_{f,t-1} + \mu_f + \mu_{s,t} + \xi_{t,h}$$

The dependent variable, $y_{f,t+h}$, is the coupon rate on firm f's newly issued LT bonds in year-quarter t + h. $EFFR_t$ is the level of EFFR. $Large_{f,t-1}$ is a dummy variable with value 1 a firm is in the topquartile of the industry-wide asset-size distribution, and 0 otherwise. $X_{f,t-1}$ is a vector of controls, including the interaction of $Large_{f,t-1}$ with several macro-controls and of $\Delta EFFR_t$ with other firm characteristics, namely lagged sales growth, leverage and liquid assets. μ_f and $\mu_{s,t}$ represent vectors of firm and industry*year-quarter fixed effects, respectively. $\xi_{t,h}$ is an error-term, double-clustered at the firm and industry*year-quarter level. The x-axis is measured in terms of quarters after the shock. The black solid line reports the point estimates for $\beta_{3,h}$; the dashed grey line the 10% confidence intervals.

Tables

TABLE 1: SUMMARY STATISTICS

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
VARIABLES	Scale	N	mean	p25	p50	p75	sd
Macro-level Variables							
$\Delta LT-Debt_{t}$	%	112	0.154	-0.360	0.0804	0.553	0.642
$\Delta LT-Debt_{t+1}$	%	111	0.315	-0.467	0.0687	0.992	1.141
$\Delta LT-Debt_{t+2}$	%	110	0.480	-0.590	0.232	1.558	1.602
$\Delta LT-Debt_{t+3}$	%	109	0.651	-0.794	0.123	1.834	2.026
$\Delta LT-Debt_{t+4}$	%	108	0.826	-0.720	0.218	2.554	2.401
$\Delta EFFR_{t}$	%	112	-0.0638	-0.105	-0.01000	0.0900	0.445
$\varepsilon^{mp.g}_{t}$	%	112	-0.0340	-0.0325	-0.00750	0.00500	0.0966
$\varepsilon^{mp.jk}_{t}$	%	112	-0.0192	-0.0396	-0.00759	0.0220	0.0756
ΔGDP_{t-1}	%	112	2.458	1.650	2.600	3.650	1.712
ΔCPI_{t-1}	%	112	2.496	1.714	2.584	3.202	1.284
Rec_{t-1}	0/1 dummy	112	0.0982	0	0	0	0.299
$\Delta LT-Treas_{t-1}$	%	112	-0.0197	-0.231	-0.00926	0.254	0.357
$\Delta i^{10y\text{-}3m}_{t-1}$	%	112	0.00866	-0.300	-0.0550	0.275	0.485
$\Delta i^{baa\text{-}aaa}_{t-1}$	%	112	-0.00205	-0.0900	-0.01000	0.0600	0.252
Firm-level Variables							
$\Delta LT-Debt_{f,t}$	%	327,532	-0.482	-2.051	0	0.800	17.03
$\Delta LT-Debt_{f,t+1}$	%	299,722	-0.915	-4.038	0	1.911	21.77
$\Delta LT-Debt_{f,t+2}$	%	284,670	-1.238	-5.662	0	2.987	24.65
$\Delta LT-Debt_{f,t+3}$	%	270,304	-1.466	-6.823	-0.00232	3.757	26.47
$\Delta LT-Debt_{f,t+4}$	%	259,167	-1.568	-7.964	0	4.795	28.08
$\Delta Sales_{f,t-1}$	%	327,532	0.75	-8.14	1.17	10.4	23
Liquid $Assets_{f,t-1}$	%	327,532	12.9	1.39	5.21	16.5	18.1
$Leverage_{f,t-1}$	%	327,532	34.8	12.00	27.4	42.9	45.7

(continues on next page)

Size$_{f,t-1}$	Log(Mln US$)	327,532	4.965	3.183	4.938	6.764	2.477
1(Issue)$_{f,t}$	0/1 dummy	118,993	0.0677	0	0	0	0.251
1(Issue)$_{f,t+1}$	0/1 dummy	110,896	0.0745	0	0	0	0.263
1(Issue)$_{f,t+2}$	0/1 dummy	106,453	0.0781	0	0	0	0.268
1(Issue)$_{f,t+3}$	0/1 dummy	102,372	0.0813	0	0	0	0.273
1(Issue)$_{f,t+4}$	0/1 dummy	99,435	0.0829	0	0	0	0.276
Coupon$_{f,t}$	%	8,046	6.119	4.060	6.125	8	2.846
Coupon$_{f,t+1}$	%	8,262	6.093	4	6.125	8	2.837
Coupon$_{f,t+2}$	%	8,309	6.080	4.016	6.125	7.920	2.824
Coupon$_{f,t+3}$	%	8,316	6.061	4	6.125	7.875	2.826
Coupon$_{f,t+4}$	%	8,240	6.030	4	6.094	7.875	2.801
Mutual Fund-level variables							
ΔMatu$_{m,t}$	%	72,389	-0.412	-3.190	-0.199	2.327	15.55
ΔMatu$_{m,t+1}$	%	72,072	-0.799	-5.241	-0.731	3.673	19.55
ΔMatu$_{m,t+2}$	%	71,617	-1.226	-6.860	-1.046	4.437	22.51
ΔMatu$_{m,t+3}$	%	70,940	-1.614	-8.246	-1.409	5.167	24.40
Mutual Fund-level variables							
ΔMatu$_{m,t+4}$	%	69,924	-1.926	-9.381	-1.816	5.555	26.17
ΔCB$_{m,t}$	%	71,063	-0.515	-4.449	-0.459	3.214	18.26
ΔCB$_{m,t+1}$	%	68,201	-1.042	-6.847	-0.974	4.485	23.23
ΔCB$_{m,t+2}$	%	65,217	-1.644	-8.635	-1.389	5.219	26.46
ΔCB$_{m,t+3}$	%	62,112	-2.111	-9.789	-1.928	5.607	29.04
ΔCB$_{m,t+4}$	%	58,889	-2.810	-10.99	-2.422	5.823	30.79
HY$_m$	0/1 dummy	72,389	0.388	0	0	1	0.487
TurnoverRatio$_{m,t-1}$	%	72,389	0.763	0.440	0.720	1.490	8.429
ExpenseRatio$_{m,t-1}$	%	72,389	0.00979	0.00570	0.00810	0.0119	0.00507
NAV$_{m,t-1}$	Log(Mln US$)	72,389	1.822	1.687	1.806	1.920	0.481
Returns$_{m,t-1}$	%	72,387	0.00960	-0.000325	0.00875	0.0215	0.0234
Returns$_{m,t-1}$	%	72,387	0.00960	-0.000325	0.00875	0.0215	0.0234

Macro-level Variables. Period: 1990-2017. $\Delta LT - Debt_{t+j}$ is the change in the aggregate LT-debt share (i.e., fraction of debt with maturity above 1-year) between year-quarter $t-1$ and year-quarter $t+j$, $j = 0, 1, ...4$. $\Delta EFFR_t$ is the quarterly variation in the Effective Funds Rate. $\varepsilon^{mp,g}$ is the 30minute surprise in FED-Funds futures around policy announcements from Gürkaynak et al. (2005) (aggregated at the quarterly frequency). $\varepsilon^{mp,jk}$ is the interest rate shock from Jarocinski & Karadi (2020). ΔGDP_{t-1} is the lagged annual GDP growth rate. ΔCPI is the lagged annual inflation rate. Rec_{t-1} is a lagged recession dummy. $\Delta LT - Treas_{t-1}$ is the lagged quarterly change in the share of Treasuries with maturity above 10-year. $\Delta i^{10y\text{-}3m}_{t-1}$ is the lagged quarterly variation of the difference between the 10-year and the 3-month yield on benchmark US Treasuries (term-spread). $\Delta i^{baa\text{-}aaa}_{t-1}$ is the lagged quarterly variation of the difference between the BAA and the AAA Moody's Seasoned Corporate Bond Yield (corporate spread).

Firm-level Variables. Period: 1990-2017. Sample: non-financial companies identified as in Ottonello & Winberry (forthcoming). $\Delta LT - Debt_{f,t+j}$ is the change in firm f's LT-debt share (i.e., fraction of debt with maturity above 1-year) between year-quarter $t - 1$ and year-quarter $t + j$, $j = 0, 1, ...4$. $\Delta Sales_{f,t-1}$ is the lagged quarterly change in log sales, expressed in p.p.. $LiquidAssets_{f,t-1}$ is the lagged share of liquid assets over total assets. $Leverage_{f,t-1}$ is the lagged ratio between total debt and total assets. $Size_{f,t-1}$ is the lagged log assets size. $\mathbf{1}(Issue_{f,t+j})$ is a dummy variable with value 1 if firm f issues bonds with maturity above 1-year in year-quarter $t + j$ and with value 0 otherwise, $j = 0, 1, ...4$. $Coupon_{f,t+j}$ is the coupon rate on the bonds issued by firm f in year-quarter $t + j$, $j = 0, 1, ...4$.

MutualFund-levelvariables. Period: 2010q2-2018q2. Sample: Corporate Bond Mutual Funds, identified as those with CRSP style categories: I, ICQH, ICQM, ICQY, ICDI, ICDS, or IC. $\Delta Matu_{m,t+4}$ is the change in the log (weighted) average portfolio maturity of fund m between year-quarter $t- 1$ and year-quarter $t + j$, $j = 0, 1, ...4$. $\Delta CB_{m,t+4}$ is the change in the log corporate bond holdings between year-quarter $t - 1$ and year-quarter $t + j$, $j = 0, 1, ...4$. HY_m is a dummy with value 1 if a fund m is classified as High-Yield, and 0 otherwise. HY funds are those with Lipper style code: HY, GB, FLX, MSI, or SFI. $TurnoverRatio_{m,t-1}$ is the lagged fund m's turnover ratio, corresponding to minimum (of aggregated sales or aggregated purchases of securities), divided by the average 12-month Total Net Assets. $ExpenseRatio_{m,t-1}$ is the lagged fund m's lagged expense ratio, i.e. the ratio of total investment that shareholders pay for the fund's operating expenses. $NAV_{m,t-1}$ is the latest (lagged) fund net asset value, i.e. the value of assets minus liabilities. $Returns_{m,t-1}$ reflects the lagged fund m's quarterly returns, computed as the growth in net asset value from one year-quarter to the next.

TABLE 2: Firm-Level Regressions

	ΔLT − Debtf,t+j																				
j =	(0)	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(16)	(17)	(18)	(19)	(20)
$Large_{f,t-1}$	0.104	-0.184	-0.333	-0.234	-0.070	-0.263	-0.576	-1.020	-1.229*	-2.048***	-2.676***	-2.945***	-2.159**	-2.954***	-3.599***	-4.471***	-4.923***	-4.808***	-4.424***	-4.177***	-4.189***
	(0.207)	(0.332)	(0.434)	(0.518)	(0.569)	(0.599)	(0.634)	(0.683)	(0.734)	(0.757)	(0.788)	(0.821)	(0.847)	(0.878)	(0.895)	(0.914)	(0.942)	(0.981)	(0.987)	(1.012)	(1.026)
$\Delta EFFR_t * Large_{f,t-1}$	0.006	-0.492**	-0.899***	-0.873**	-1.376***	-1.601***	-2.151***	-1.538***	-1.799***	-1.716***	-1.571***	-1.120**	-0.304	-0.158	-0.164	-0.006	-0.026	-0.177	-0.048	-0.095	-0.018
	(0.163)	(0.242)	(0.306)	(0.368)	(0.399)	(0.421)	(0.435)	(0.454)	(0.470)	(0.484)	(0.489)	(0.498)	(0.508)	(0.524)	(0.536)	(0.547)	(0.547)	(0.549)	(0.550)	(0.538)	(0.544)
$\Delta Sales_{f,t-1}$	0.013***	0.024***	0.005*	0.006**	0.018***	0.021***	0.004	0.005	0.014***	0.018***	0.003	0.003	0.017***	0.021***	0.002	0.001	0.009**	0.012***	-0.003	-0.000	0.016***
	(0.002)	(0.002)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.005)
$\Delta EFFR_t * \Delta Sales_{f,t-1}$	-0.000	-0.000	-0.004	0.003	0.003	0.006	0.003	-0.001	0.004	0.001	0.005	0.005	0.000	0.014	-0.002	0.004	0.007	0.008	0.009	-0.005	-0.006
	(0.004)	(0.005)	(0.006)	(0.006)	(0.007)	(0.007)	(0.007)	(0.007)	(0.008)	(0.008)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
Liquid $Assets_{f,t-1}$	-0.017***	-0.038***	-0.053***	-0.067***	-0.079***	-0.083***	-0.080***	-0.079***	-0.084***	-0.094***	-0.096***	-0.083***	-0.089***	-0.093***	-0.099***	-0.091***	-0.086***	-0.087***	-0.078***	-0.069***	-0.084***
	(0.004)	(0.006)	(0.008)	(0.010)	(0.012)	(0.013)	(0.014)	(0.015)	(0.016)	(0.017)	(0.018)	(0.019)	(0.020)	(0.021)	(0.022)	(0.023)	(0.024)	(0.024)	(0.024)	(0.025)	(0.026)
$\Delta EFFR_t *$ Liquid $Assets_{f,t-1}$	0.002	0.001	0.003	0.019	0.029*	0.020	0.023	0.021	0.025	0.026	0.019	0.008	0.007	0.010	0.011	0.016	0.002	0.005	0.027	0.042*	0.035
	(0.005)	(0.008)	(0.011)	(0.014)	(0.015)	(0.016)	(0.018)	(0.018)	(0.019)	(0.020)	(0.021)	(0.021)	(0.021)	(0.021)	(0.021)	(0.021)	(0.021)	(0.022)	(0.022)	(0.022)	(0.023)
$Leverage_{f,t-1}$	-0.000	-0.001	-0.002	-0.004	-0.007	-0.007	-0.008	-0.010	-0.008	-0.011	-0.014*	-0.017*	-0.019**	-0.026***	-0.031***	-0.032***	-0.031***	-0.035***	-0.040***	-0.045***	-0.047***
	(0.001)	(0.002)	(0.003)	(0.004)	(0.005)	(0.006)	(0.006)	(0.007)	(0.007)	(0.008)	(0.008)	(0.009)	(0.009)	(0.009)	(0.010)	(0.010)	(0.010)	(0.010)	(0.011)	(0.011)	(0.011)
$\Delta EFFR_t * Leverage_{f,t-1}$	-0.003*	-0.002	-0.003	-0.000	-0.002	-0.001	0.002	-0.006	-0.006	-0.013	-0.010	-0.015*	-0.021**	-0.015*	-0.015*	-0.009	-0.012	-0.015*	-0.011	-0.010	-0.016**
	(0.002)	(0.003)	(0.003)	(0.004)	(0.005)	(0.006)	(0.006)	(0.007)	(0.007)	(0.008)	(0.008)	(0.008)	(0.009)	(0.009)	(0.009)	(0.009)	(0.008)	(0.009)	(0.008)	(0.009)	(0.008)
Observations	327,532	299,430	284,233	269,838	258,377	248,691	241,564	232,489	224,179	216,403	210,684	203,171	196,431	190,124	185,563	179,253	173,641	168,203	164,147	158,581	153,430
R-squared	0.093	0.119	0.138	0.154	0.170	0.183	0.192	0.202	0.212	0.223	0.231	0.240	0.248	0.260	0.266	0.276	0.281	0.292	0.300	0.311	0.320

In column j, the dependent variable is $\Delta LT - Debt_{f,t+j}$, i.e., the change in firm f's share of LT-debt between year-quarter $t-1$ and $t+j$. $\Delta EFFR_t$ is the quarterly variation in the Effective FED Funds Rate. $Large_{f,t-1}$ is a dummy with value 1 if firm f is in the top asset-size quartile of the respective (3-digit SIC) industry distribution. $\Delta Sales_{f,t-1}$ is the quarterly variation in firm f's log sales. Liquid $Assets_{f,t-1}$ is the share of liquid assets by firm f. $Leverage_{f,t-1}$ is firm f's leverage, defined as total debt over total assets. All firm-level variables are lagged by one year-quarter. Each regression additionally includes the full interaction of $Large_{f,t-1}$ with a vector of lagged macro controls (annual GDP growth and inflation rate; a recession dummy; quarterly variation in term-spread, corporate spread and in share of Treasuries with maturity above 20-year). Furthermore, in each column we apply firm and industry*year-quarter fixed effects. Standard errors are clustered at the Firm and Industry*Year-Quarter level. *** p<0.01, ** p<0.05, * p<0.1.

TABLE 3: Monetary Policy and Likelihood of Issuing LT Bonds – Relative Response of Large Companies

	(1)	(2)	(3)	(4)	(5)	(6)
		$1(\text{Issue})_{f,t}$			$1(\text{Issue})_{f,t+1}$	
$\Delta EFFR_t$	-0.010***			-0.012***		
	(0.003)			(0.004)		
$Large_{f,t-1}$	0.047***	0.046***	0.047***	0.043***	0.044***	0.039***
	(0.005)	(0.006)	(0.006)	(0.005)	(0.006)	(0.007)
$\Delta EFFR_t * Large_{f,t-1}$	-0.013***	-0.008**	-0.011**	-0.018***	-0.013***	-0.016***
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.005)
Observations	118,993	118,993	112,669	111,703	111,703	105,198
R-squared	0.089	0.094	0.210	0.091	0.096	0.218
Firm Controsl*$\Delta EFFR_t$	Yes	Yes	Yes	Yes	Yes	Yes
Macro Controls*Large	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Year:Quarter FE	No	Yes	-	No	Yes	-
Industry*Year-Quarter FE	No	No	Yes	No	No	Yes

In columns (1)-(3), the dependent is variable, $1(\text{Issue})_{f,t}$ is a dummy variable with value if firm f issues LT bonds in year-quarter t. In columns (3)-(6), the dependent is the same variable, though measuread as of year-quarter t + 1. $\Delta EFFR_t$ is the quarterly variation in the Effective FED Funds Rate. $Large_{f,t-1}$ is a dummy with value 1 if firm f is in the top asset-size quartile of the respective (3-digit SIC) industry distribution. Firm controls include lagged sales growth, leverage and share of liquid assets. Macro controls are given by annual GDP growth and inflation rate; a recession dummy; quarterly variation in term-spread, corporate spread and in share of Treasuries with maturity above 20-year. Standard errors are clustered at the Firm and Industry*Year-Quarter level. *** p<0.01, ** p<0.05, * p<0.1.

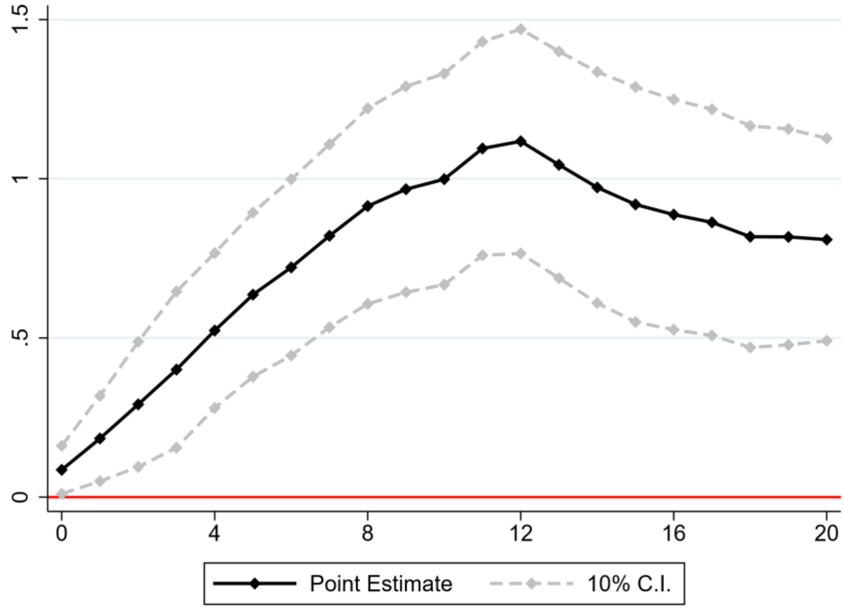
TABLE 4: Monetary Policy and Financing Costs Through LT Bonds

	(1)	(2)	(3)	(4)
	$Coupon_{f,t}$		$Coupon_{f,t+1}$	
$EFFR_t$	0.092**		0.133***	
	(0.043)		(0.045)	
$Large_{f,t-1}$	-1.573***	-0.065	-1.516***	-0.181
	(0.340)	(0.473)	(0.296)	(0.537)
$EFFR_t * Large_{f,t-1}$	0.269***	0.118**	0.231***	0.108**
	(0.041)	(0.052)	(0.039)	(0.054)
Observations	7,310	4,157	7,551	4,215
R-squared	0.706	0.835	0.711	0.836
Firm Controls*EFFR	Yes	Yes	Yes	Yes
Macro Controls*Large	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes
Industry*Year:Quarter FE	No	Yes	No	Yes

In columns (1)-(2), the dependent variable, is the coupon rate on firm f's newly issued LT bonds in year-quarter t. In columns (3)-(4), the left-hand side variable is the same, though measured in yearquarter t + 1. $EFFR_t$ is the level of EFFR. $Large_{f,t-1}$ is a dummy with value 1 if firm f is in the top asset-size quartile of the respective (3-digit SIC) industry distribution. Firm controls include lagged sales growth, leverage and share of liquid assets. Macro controls are given by annual GDP growth and inflation rate; a recession dummy; term-spread, corporate spread and share of Treasuries with maturity above 20-year. Standard errors are double-clustered at the firm and industry*year-quarter level. *** p<0.01, ** p<0.05, * p<0.1.

Empirical Appendix

FIGURE A1: Monetary Policy and Debt Maturity Structure: Aggregate Response - Pre-Crisis Period

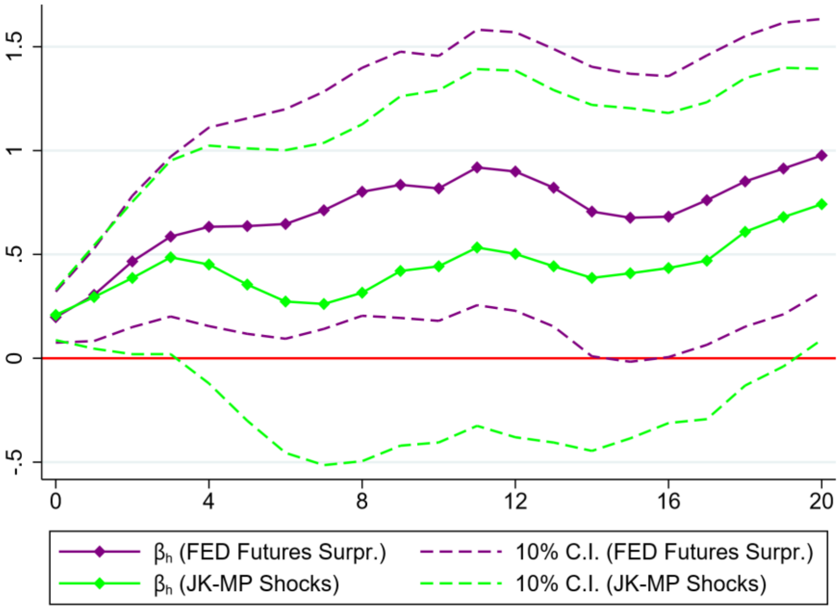


This figure depicts the response of the aggregate-level share of LT debt to a 25 b.p. cut in the EFFR. The sample includes observations from 1990q1 to 2008q4. Formally, it shows the coefficients $\beta_{1,h}$ from the estimation of the following local projection model:

$$\Delta_h y_{t+h} = \beta_{1,h} \Delta EFFR_t + MacroControls_{t-1} + u_{t,h}$$

The dependent variable, $\Delta_h y_{t+h}$, represents the growth of the LT-debt share (expressed in p.p.) from year-quarter $t - 1$ to year-quarter $t + h$. $\Delta EFFR_t$ is the quarterly EFFR change. $MacroControls_{t-1}$ is a vector of lagged macroeconomic controls, including annual GDP growth and inflation rate, a dummy for recessions, the quarterly variation in the share of LT treasuries, in the corporate spread and in the 10y-3m term-spread. $u_{t,h}$ is a robust error-term. The x-axis is measured in terms of quarters after the shock. The black solid line reports the point estimates for $\beta_{1,h}$; the dashed grey line the 10% confidence intervals.

FIGURE A2: Monetary Policy and Debt Maturity Structure: Aggregate Response using Exogenous Shocks - Pre-Crisis Period

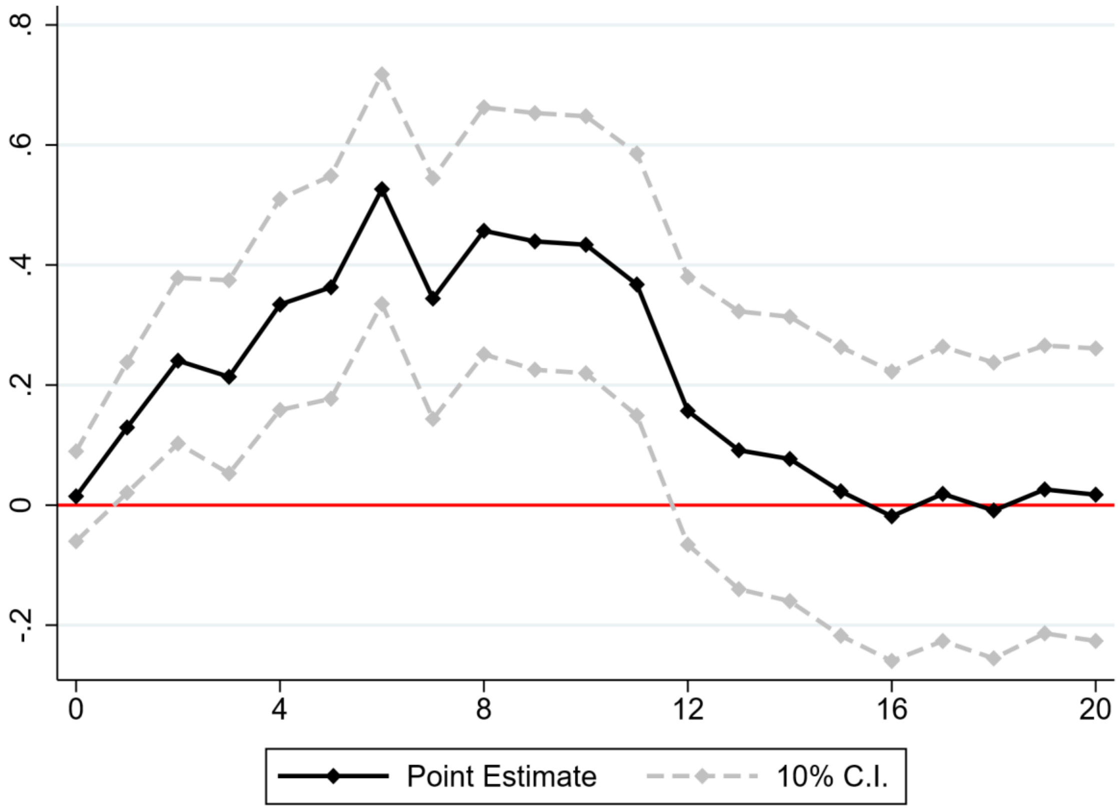


This figure depicts the response of the aggregate-level share of LT debt to a 1 s.d. reduction in the monetary policy shock. The sample includes observations from 1990q1 to 2008q4. Formally, it shows the coefficients $\beta_{1,h}$ from the estimation of the following local projection model:

$$\Delta_h y_{t+h} = \beta_{1,h}\varepsilon^{mp}{}_t + \text{MacroControls}_{t-1} + u_{t,h}$$

The dependent variable, $\Delta_h y_{t+h}$, represents the growth of the LT-debt share (expressed in p.p.) from year-quarter $t - 1$ to year-quarter $t + h$. $\varepsilon^{mp}{}_t$ is an exogenous monetary policy shock, gathered from either Gürkaynak et al. (2005) or from Jarocinski & Karadi (2020). $\text{MacroControls}_{t-1}$ is a vector of lagged macroeconomic controls, including annual GDP growth and inflation rate, a dummy for recessions, the quarterly variation in the share of LT treasuries, in the corporate spread and in the 10y-3m term-spread. $u_{t,h}$ is a robust error-term. The x-axis is measured in terms of quarters after the shock. The purple (green) solid line, connected by diamonds (squares), reports the point estimates for $\beta_{1,h}$ using the Gürkaynak et al. 2005 (Jarocinski & Karadi 2020) shocks and the dashed (dotted) purple (green) line the respective 10% confidence intervals.

FIGURE A3: Monetary Policy and Debt Maturity Structure - Relative Response of Large Companies: Pre-Crisis Period

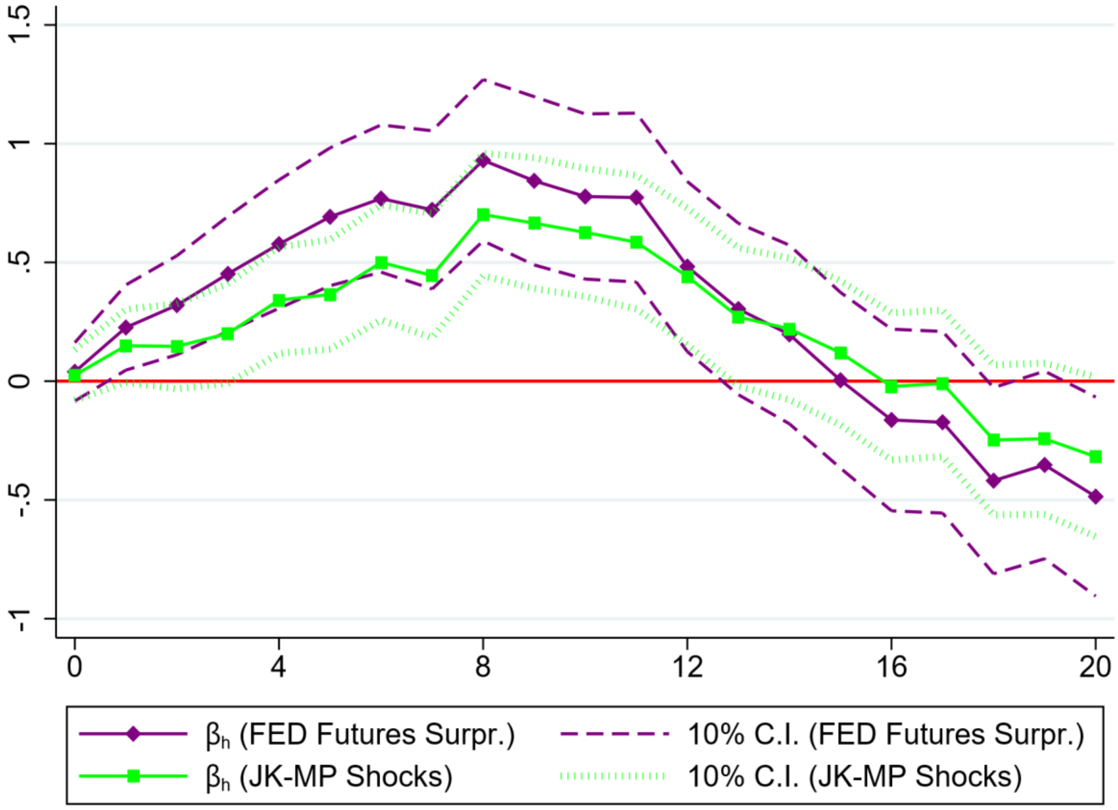


This figure depicts the relative response of companies in the top quartile of the 3-digit SIC industry asset-size distribution to a 25 b.p. cut in the EFFR (as compared to smaller firms). The regression sample goes from 1990q1 to 2008q4. Formally, the picture shows the coefficients $\beta_{3,h}$ from the estimation of the following local projection model:

$$\Delta_h y_{f,t+h} = \beta_{1,h}\Delta EFFR_t + \beta_{2,h}Large_{f,t-1} + \beta_{3,h}Large_{f,t-1} * \Delta EFFR_t + X_{f,t-1} + \mu_f + \mu_{s,t} + u_{t,h}$$

The dependent variable, $\Delta_h y_{f,t+h}$, represents the growth of the share of LT-debt - expressed in p.p. from year-quarter $t-1$ to year-quarter $t+h$. $\Delta EFFR_t$ is the quarterly change in the EFFR. $Large_{f,t-1}$ is a dummy variable with value 1 a firm is in the top-quartile of the industry-wide asset-size distribution, and 0 otherwise. $X_{f,t-1}$ is a vector of controls, including the interaction of $Large_{f,t-1}$ with several macro-controls and of $\Delta EFFR_t$ with other firm characteristics, namely lagged sales growth, leverage and liquid assets. μ_f and $\mu_{s,t}$ represent vectors of firm and industry*year-quarter fixed effects, respectively. $u_{t,h}$ is an error-term, double-clustered at the firm and industry*year-quarter level. The x-axis is measured in terms of quarters after the shock. The black solid line reports the point estimates for $\beta_{3,h}$; the dashed grey line the 10% confidence intervals.

FIGURE A4: Monetary Policy and Debt Maturity Structure - Relative Response of Large Firms using Exogenous Shocks: Pre-Crisis

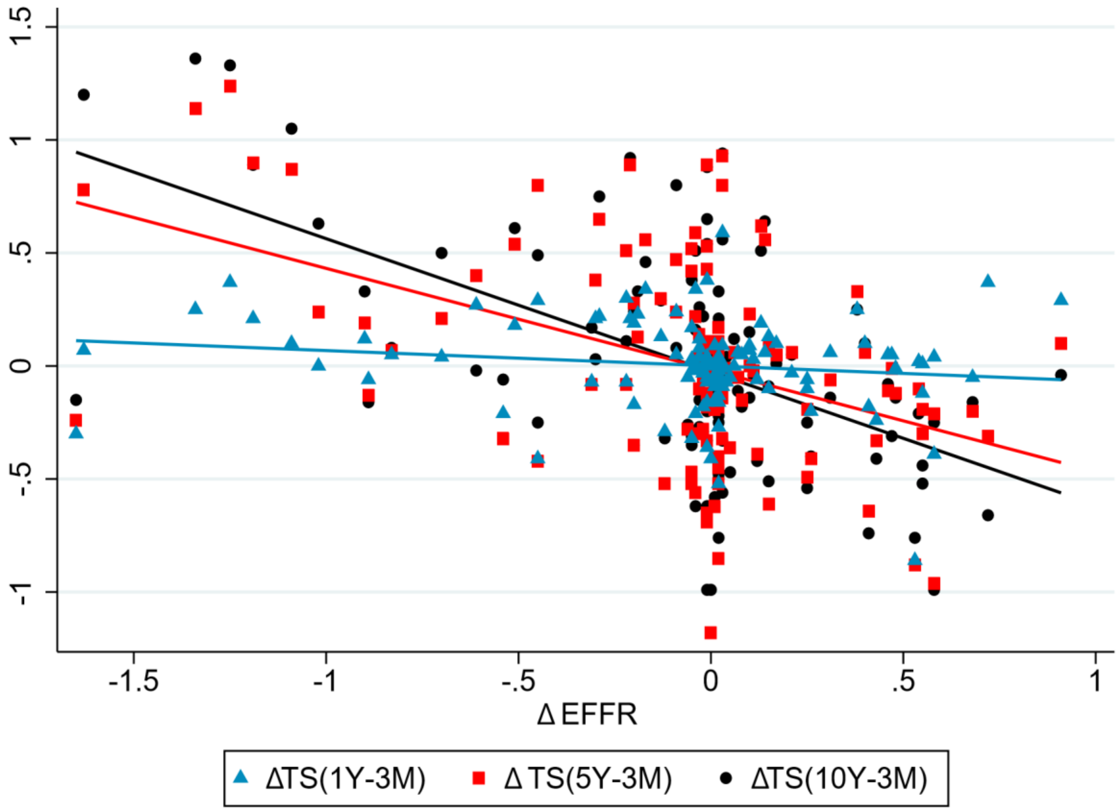


This figure depicts the relative response of companies in the top quartile of the 3-digit SIC industry asset-size distribution to a 1 s.d. b.p.reduction in monetary policy shock (as compared to smaller firms). The regression sample goes from 1990q1 to 2008q4. Formally, the picture shows the coefficients $\beta_{3,h}$ from the estimation of the following local projection model:

$$\Delta_h y_{f,t+h} = \beta_{1,h}\varepsilon^{mp}_t + \beta_{2,h}Large_{f,t-1} + \beta_{3,h}Large_{f,t-1} * \varepsilon^{mp}_t + X_{f,t-1} + \mu_f + \mu_{s,t} + u_{t,h}$$

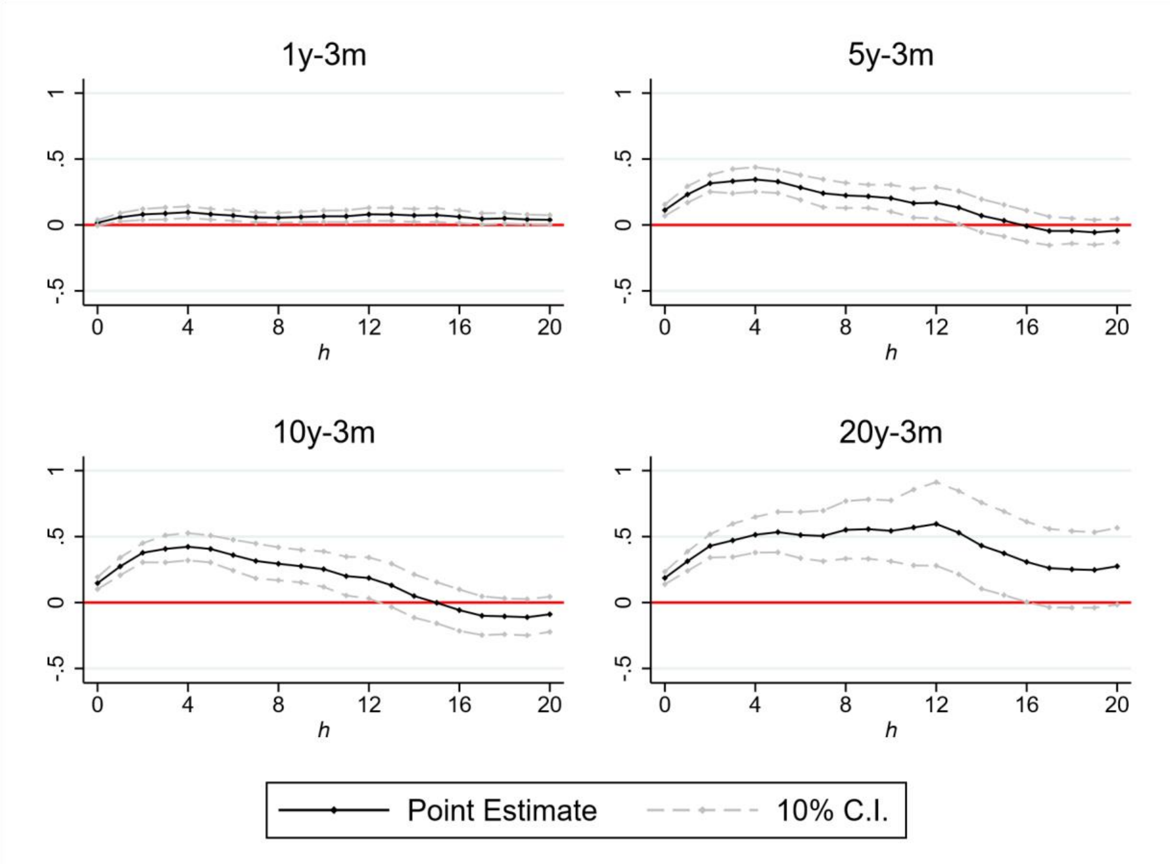
The dependent variable, $\Delta_h y_{f,t+h}$, represents the growth of the share of LT-debt - expressed in p.p. from year-quarter $t-1$ to year-quarter $t+h$. ε^{mp}_t is an exogenous monetary policy shock, derived either from Gürkaynak et al. (2005) or from Jarocinski & Karadi (2020). $Large_{f,t-1}$ is a dummy variable with value 1 a firm is in the top-quartile of the industry-wide asset-size distribution, and 0 otherwise. $X_{f,t-1}$ is a vector of controls, including the interaction of $Large_{f,t-1}$ with several macro-controls and of ε^{mp}_t with other firm characteristics, namely lagged sales growth, leverage and liquid assets. μ_f and $\mu_{s,t}$ represent vectors of firm and industry*year-quarter fixed effects, respectively. $u_{t,h}$ is an errorterm, double-clustered at the firm and industry*year-quarter level. The x-axis is measured in terms of quarters after the shock. The purple (green) solid line, connected by diamonds (squares), reports the point estimates for $\beta_{1,h}$ using the Gürkaynak et al. 2005 (Jarocinski & Karadi 2020) shocks and the dashed (dotted) purple (green) line the respective 10% confidence intervals.

FIGURE A5: Contemporaneous Quarterly Variation of Term-Spread and EFFR



We report the quarterly variation of the Effective FED Funds Rate on the x-axis and the contemporaneous quarterly change in the term-spread on the y-axis. Both measures are expressed in percentage points. We employ several definitions of the term-spread, all based on the benchmark Treasury yields at constant maturity. The 1-year/3-month spread is in light blue. The 5-year/3-month spread is in red, whereas the 10-year/3-month spread is in black. The lines, which are colored accordingly, reflect results from simple bivariate linear-fit regressions.

FIGURE A6: Monetary Policy and the Term-Spread - Local Projections

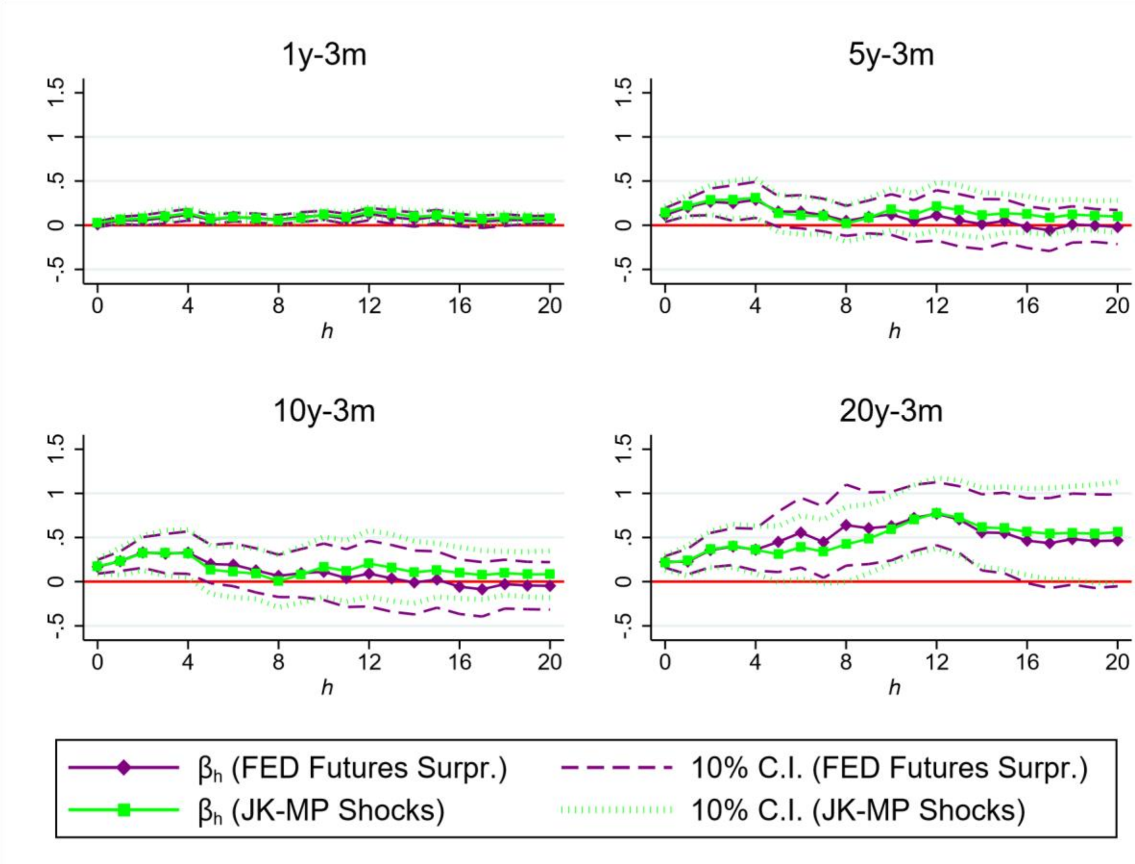


This figure depicts the response of the term-spread to a 25 b.p. cut in the EFFR. We employ several definitions of the term-spread, all based on the benchmark Treasury yields at constant maturity. The 1-year/3-month spread is in the north-west sub-plot. The 5-year/3-month spread is in the north-east sub-plot. The 10-year/3-month spread is in the south-west sub-plot, whereas the 20-year/3-month spread is in the south-east one. Formally, it shows the coefficients $\beta_{1,h}$ from the estimation of the following local projection model:

$$\Delta_h y_{t+h} = \beta_{1,h} \Delta EFFR_t + MacroControls_{t-1} + u_{t,h}$$

The dependent variable, $\Delta_h y_{t+h}$, represents the growth of the term-spread (expressed in p.p.) from year-quarter $t - 1$ to year-quarter $t + h$. $\Delta EFFR_t$ is the quarterly EFFR change. $MacroControls_{t-1}$ is a vector of lagged macroeconomic controls, including annual GDP growth and inflation rate, a dummy for recessions, the quarterly variation in the share of LT treasuries and in the corporate spread. $u_{t,h}$ is a robust error-term. The x-axis is measured in terms of quarters after the shock. The black solid line reports the point estimates for $\beta_{1,h}$; the dashed grey line the 10% confidence intervals.

FIGURE A7: Monetary Policy and the Term-Spread - Local Projections using Exogenous Shocks

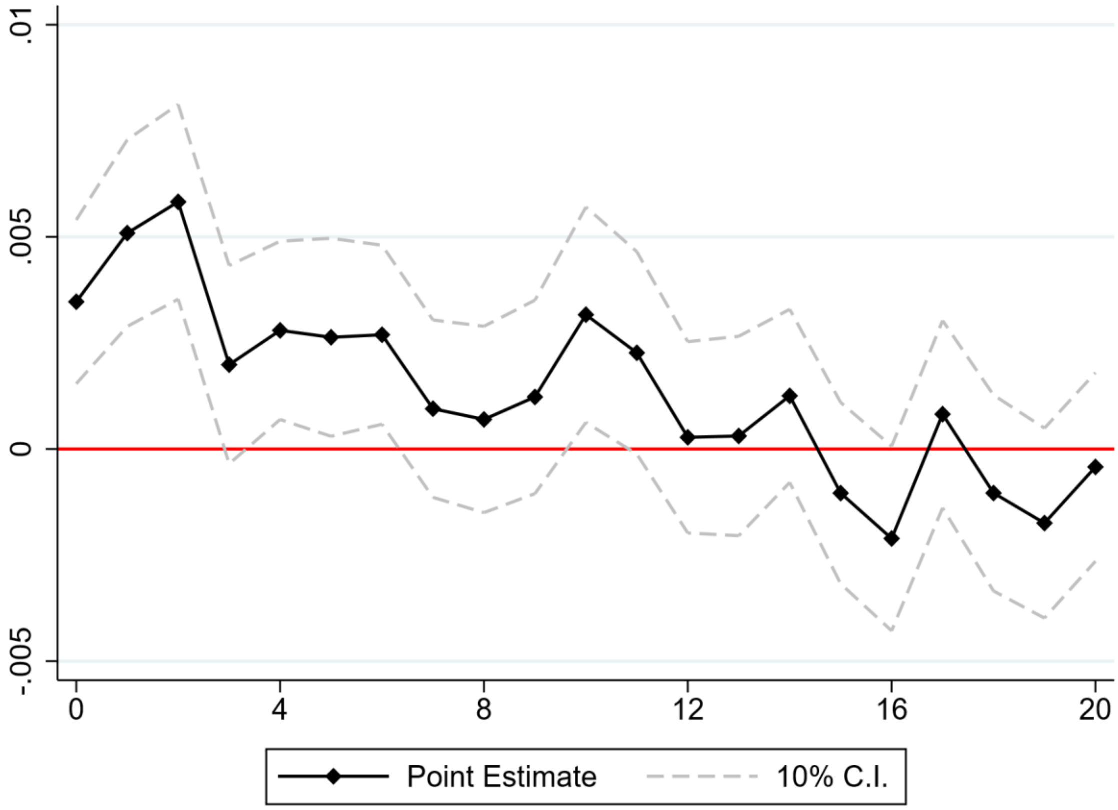


This figure depicts the response of the term-spread to a 1 s.d. reduction in the monetary policy shock. We employ several definitions of the term-spread, all based on the benchmark Treasury yields at constant maturity. The 1-year/3-month spread is in the north-west sub-plot. The 5-year/3-month spread is in the north-east sub-plot. The 10-year/3-month spread is in the south-west sub-plot, whereas the 20-year/3-month spread is in the south-east one. Formally, it shows the coefficients $\beta_{1,h}$ from the estimation of the following local projection model:

$$\Delta_h y_{t+h} = \beta_{1,h}\varepsilon^{mp}_t + MacroControls_{t-1} + u_{t,h}$$

The dependent variable, $\Delta_h y_{t+h}$, represents the growth of the term-spread (expressed in p.p.) from year-quarter t − 1 to year-quarter t + h. ε^{mp}_t is an exogenous monetary policy shock, gathered from either Gürkaynak et al. (2005) or from Jarocinski & Karadi (2020). $MacroControls_{t-1}$ is a vector of lagged macroeconomic controls, including annual GDP growth and inflation rate, a dummy for recessions, the quarterly variation in the share of LT treasuries and in the corporate spread. $u_{t,h}$ is a robust error-term. The x-axis is measured in terms of quarters after the shock. The purple (green) solid line, connected by diamonds (squares), reports the point estimates for $\beta_{1,h}$ using the Gürkaynak et al. 2005 (Jarocinski & Karadi 2020) shocks and the dashed (dotted) purple (green) line the respective 10% confidence intervals.

FIGURE A8: Monetary Policy and Likelihood of Issuing LT Bonds Relative Response of Large Companies: Pre-Crisis

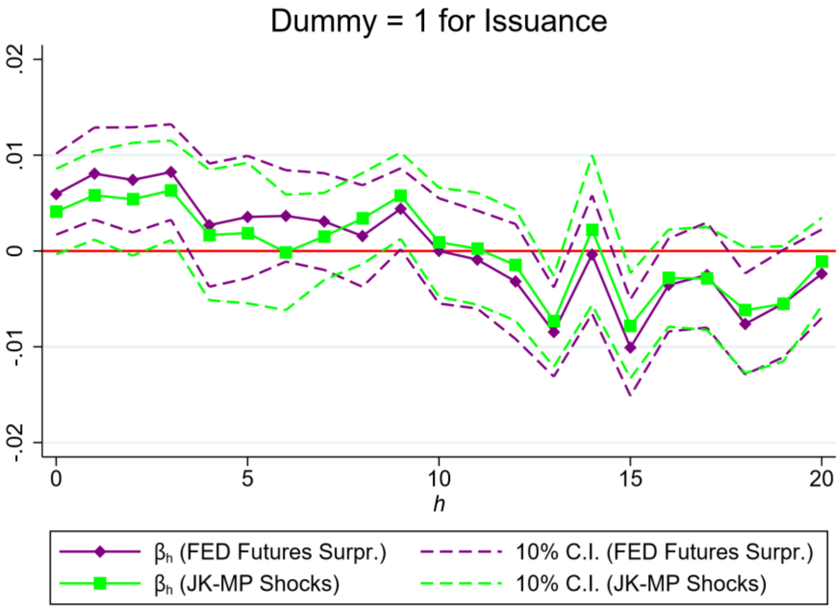


This figure depicts the relative response of companies in the top quartile of the 3-digit SIC industry asset-size distribution to a 25 b.p. cut in the EFFR (as compared to smaller firms). The regression sample goes from 1990q1 to 2008q4. Formally, it shows the coefficients $\beta_{3,h}$ from the estimation of the following local projection model:

$$y_{f,t+h} = \beta_{1,h}\Delta EFFR_t + \beta_{2,h}Large_{f,t-1} + \beta_{3,h}Large_{f,t-1} * \Delta EFFR_t + X_{f,t-1} + \mu_f + \mu_{s,t} + u_{t,h}$$

The dependent variable, $y_{f,t+h}$, is a dummy with value 1 if firms f issues LT bonds in year-quarter t + h and with value 0 if it does not. $\Delta EFFR_t$ is the quarterly change in the EFFR. $Large_{f,t-1}$ is a dummy variable with value 1 a firm is in the top-quartile of the industry-wide asset-size distribution, and 0 otherwise. $X_{f,t-1}$ is a vector of controls, including the interaction of $Large_{f,t-1}$ with several macro-controls and of $\Delta EFFR_t$ with other firm characteristics, namely lagged sales growth, leverage and liquid assets. μ_f and $\mu_{s,t}$ represent vectors of firm and industry*year-quarter fixed effects, respectively. $u_{t,h}$ is an error-term, double-clustered at the firm and industry*year-quarter level. The x-axis is measured in terms of quarters after the shock. The black solid line reports the point estimates for $\beta_{3,h}$; the dashed grey line the 10% confidence intervals.

FIGURE A9: Monetary Policy and Likelihood of Issuing LT Bonds Relative Response of Large Companies using Exogenous Shocks: Pre-Crisis



This figure depicts the relative response of companies in the top quartile of the 3-digit SIC industry asset-size distribution to a 1 s.d. reduction in the monetary policy shock. (as compared to smaller firms). Formally, it shows the coefficients $\beta_{3,h}$ from the estimation of the following local projection model:

$$y_{f,t+h} = \beta_{1,h}\varepsilon^{mp}_t + \beta_{2,h}Large_{f,t-1} + \beta_{3,h}Large_{f,t-1} * \varepsilon^{mp}_t + X_{f,t-1} + \mu_f + \mu_{s,t} + u_{t,h}$$

The dependent variable, $y_{f,t+h}$, is a dummy with value 1 if firms f issues LT bonds in year-quarter t+h and with value 0 if it does not. ε^{mp}_t is an exogenous monetary policy shock, gathered from either Gürkaynak et al. (2005) or from Jarocinski & Karadi (2020). $MacroControls_{t-1}$ is a vector of lagged macroeconomic controls, including annual GDP growth and inflation rate, a dummy for recessions, the quarterly variation in the share of LT treasuries, in the corporate spread and in the 10y-3m termspread. $X_{f,t-1}$ is a vector of firm-level controls, including lagged sales growth, leverage and liquid assets. μ_f is a vector of firm fixed effects. $u_{t,h}$ is an error-term, double-clustered at the firm and industry*year-quarter level. The x-axis is measured in terms of quarters after the shock. The purple (green) solid line, connected by diamonds (squares), reports the point estimates for $\beta_{1,h}$ using the Gürkaynak et al. 2005 (Jarocinski & Karadi 2020) shocks and the dashed (dotted) purple (green) line the respective 10% confidence intervals.

Theory Appendix

PROOFS

Proof of Lemma 1. Firms' endowments are distributed uniformly and continuously on [0,I]. This implies that each firm is atomistic relative to the set of all firms. The expected return of the market portfolio thus does not respond to changes in a single firm's $d_l(A)$ and $\rho^*(A)$:

$$\frac{\partial \int_0^I P_l(A)d_l(A)dA}{\partial d_l(A)}=0 \qquad \frac{\partial \int_0^I P_l(A)d_l(A)dA}{\partial \rho^*(A)}=0$$

Taking derivative of (11) yields:

$$\frac{\partial P_l(A)}{\partial d_l(A)}=0 \qquad \frac{\partial P_l(A)}{\partial \rho^*(A)}=\delta f(\rho^*(A))\frac{1-\gamma(1-\delta)\int_0^I P_l(A)d_l(A)dA}{(1+i_1)[\alpha(1+i_1)+(1-\alpha)(1+i_2)]}$$

Note that: $\frac{\partial P_l(A)}{\partial \rho^*(A)}>0$ iff $P_l(A)>0$, which holds given the firm first order condition in equation (7).

□

Proof of Proposition 1. The three constraints of the firms' problem are:

$$r-\rho^*(A)\geq d_s(A) \qquad \text{(LL)}$$

$$R-\frac{B}{\Delta P}\geq d_l(A) \qquad \text{(IC)}$$

$$\frac{d_s(A)}{1+i_1}+P_l(A)d_l(A)\geq I-A \qquad \text{(IR)}$$

First, we will show that these three constraints will only bind simultaneously, such that, we refer to one Lagrange-multiplier $\lambda(A)$. If there exists any one firm that is unconstrained in equilibrium, i.e., if there exists A' such that $0\leq A'\leq I$ and $\lambda_2(A')=\lambda_3(A')=0$, then (11) implies that:

$$P_l(A')=\frac{1+\lambda_2(A')}{1+\lambda_3(A')}\frac{\delta F(\rho^*(A'))}{(1+i_1)(1+i_2)}=\frac{\delta F(\rho^*(A'))}{(1+i_1)(1+i_2)}.$$

Since it is true (as (7) must hold for all possible endowments A) that

$$\frac{P_l(A)}{\delta F(\rho^*(A))}=\frac{1+\lambda_2(A)}{1+\lambda_3(A)}\frac{1}{(1+i_1)(1+i_2)}$$

and since the LHS is identical for all A, it follows that

$$\frac{1+\lambda_2(A)}{1+\lambda_3(A)}=1 \;\Leftrightarrow \lambda_3(A)=\lambda_3(A) \;\; \forall A$$

and from (10) that

$$\lambda(A)\equiv\lambda_1(A)=\lambda_2(A)=\lambda_3(A) \quad \forall A.$$

Now, we can pinpoint the threshold endowment $\overline{A}$, which is is the lowest endowment at which $\lambda(\overline{A})=0$. The critical value $\overline{A}$ is the one that makes the constraints bind exactly, for a firm that chooses the optimal cutoff value $\rho^*=\delta R/(1+i_2)$. We find it by plugging the LL and the IC constraints into the IR constraint and from the FOC with respect to $d_l(A)$., the price of LT debt takes the form

$$P_l(A)=\frac{\delta F\left(\frac{\delta R}{1+i_2}\right)}{(1+i_1)(1+i_2)}$$

and we get that the threshold endowment must be:

$$\overline{A}=I-\frac{r-}{1+i_1}-\frac{\delta F\left(\frac{\delta R}{1+i_2}\right)}{(1+i_1)(1+i_2)}\left(R-\frac{B}{\Delta P}\right)$$

Both types, unconstrained and constrained firms, exist concurrently if $\overline{A} \in (0, I)$.

□

Proof of Lemma 2. Combining (11) and (12):

$$\rho^*(A)=\frac{\delta R}{1+i_2}-\lambda(A)\left[\frac{1}{f(\rho^*)}-\frac{\delta d_l(A)}{1+i_2}\right].$$

An increase in $\rho^*(A)$ has two effects for constrained firms. On the one hand, it tightens the LL constraint, such that less short-term debt can be issued. On the other hand, a higher probability of continuation yields a higher price. We will show that our assumption $\chi\delta(1 + i_2)^{-1} (R - B/\Delta p) < 1$ assures that the former effect always dominates the latter and that all statements in this lemma follow from this fact. Note that for constrained firms, for which $\lambda(A) > 0$, it is true that:

$$\rho^*(A)<\frac{\delta R}{1+i_2} \quad \Leftrightarrow \quad f(\rho^*)\frac{\delta d_l(A)}{1+i_2}<1.$$

The last inequality $f\left(\rho^*(A)\right)\frac{\delta d_l(A)}{1+i_2}<1$ holds due to the assumption that $\chi\delta(1 + i_2)^{-1} (R - B/\Delta p) < 1$, as

$$f(\rho^*)\frac{\delta d_l(A)}{1+i_2}<f(0)\frac{\delta\left(R-\frac{B}{\Delta P}\right)}{1+i_2}<\frac{\chi\delta}{1+i_2}\ (R - B/\Delta p) < 1.$$

Therefore,

$$\rho^*(A)<\frac{\delta R}{1+i_2}.$$

Statements 1 and 2 in the Lemma follow immediately from the IR constraint. As the unconstrained firms, those with $A >\overline{A}$, always choose $\rho^*(A) = \delta R/(1 + i_2)$, there will be no effect of a change in either A nor i_1 on their choice of $\rho^*(A)$. For the constrained firms we can derive the change in $\rho^*(A)$ from the constraints.

Recall that for the constrained firm the following must hold:

$$\frac{r-\rho^*(A)}{1+i_1}+\delta F\left(\rho^*(A)\right)\frac{1-\gamma(1-\delta)\left(\int_0^I F\left(\rho^*(A)\right)d_l(A)dA-g\right)}{(1+i_1)[\alpha(1+i_1)+(1-\alpha)(1+i_2)]}\ (R - B/\Delta p)=I-A.$$

By plugging in the equilibrium price $P_l(A)=\frac{\delta F\left(\rho^*(A)\right)}{(1+i_1)(1+i_2)}$, the last expression reads as

$$\text{(l)} \quad \frac{r-\rho^*(A)}{1+i_1}+\frac{\delta F\left(\rho^*(A)\right)}{(1+i_1)(1+i_2)}-I+A=0.$$

Applying the Implicit Function Theorem to the equality above, we obtain for $A<\overline{A}$

$$\frac{\partial\rho^*}{\partial i_1}=-\frac{\frac{\partial l}{\partial i_1}}{\frac{\partial l}{\partial\rho^*(A)}}<0.$$

Now, since $\frac{\partial l}{\partial i_1}<0$ and

$$\frac{\partial l}{\partial \rho^*(A)} = -\frac{1}{1+i_1} + \frac{\delta f\left(\rho^*(A)\right)}{(1+i_1)(1+i_2)}(R - B/\Delta p) < 0$$

as

$$\frac{\delta f\left(\rho^*\right)}{1+i_2}(R - B/\Delta p) < 1.$$

Moreover,

$$\frac{\partial \rho^*(A)}{\partial A} = -\frac{1}{\frac{\partial l}{\partial \rho^*(A)}} > 0$$

□

Proof of Proposition 2. Since the unconstrained firms, i.e., those with $A > \bar{A}$, always choose $\rho^*(A) = \delta R/(1+i_2)$, there will be no effect of a change in i_1 on their choice. Furthermore, the total expected revenue from long term debt $\int_A F\left(\rho^*(A)\right) d_l(A)dA$ is pinned down by the market clearing equation:

$$\frac{1-\gamma(1-\delta)\left(\int_0^I F\left(\rho^*(A)\right) d_l(A)dA - g\right)}{[\alpha(1+i_1)+(1-\alpha)(1+i_2)]} = \frac{1}{1+i_2}$$

We can reformulate this equation as

$$\int_0^I F\left(\rho^*(A)\right) d_l(A)dA = g + \frac{\chi(i_2-i_1)}{(1-\delta)(1+i_2)}$$

We can then split the left-hand side into the unconstrained and constrained components of LT debt

$$\int_0^{\bar{A}} F\left(\rho^*(A)\right) d_l(A)dA + \int_{\bar{A}}^I F\left(\rho^*(A)\right) d_l(A)dA = g + \frac{\chi(i_2-i_1)}{(1-\delta)(1+i_2)}$$

to get

$$\int_{\bar{A}}^I d_l(A)dA = \frac{1}{F\left(\frac{\delta R}{1+i_2}\right)}\left[g + \frac{\chi(i_2-i_1)}{(1-\delta)(1+i_2)} - (R - B/\Delta p)\int_0^{\bar{A}} F\left(\rho^*(A)\right) dA\right]$$

Thus,

$$\frac{\partial \int_{\bar{A}}^I d_l(A)dA}{\partial i_1} = \frac{1}{F\left(\frac{\delta R}{1+i_2}\right)}\left[\frac{-\chi}{(1-\delta)(1+i_2)} - (R - B/\Delta p)\frac{\partial \int_0^{\bar{A}} F\left(\rho^*(A)\right) dA}{\partial i_1}\right]$$

The last derivative is negative if:

$$\frac{-\chi}{(1-\delta)(1+i_2)} - (R - B/\Delta p)\frac{\partial \int_0^{\bar{A}} F\left(\rho^*(A)\right) dA}{\partial i_1} < 0.$$

Plugging in the explicit expressions into this inequality, we have:

$$\frac{\chi}{(1-\delta)(1+i_2)} > (R - B/\Delta p)\int_0^{\bar{A}}\left(f\left(\rho^*(A)\right)\frac{\frac{r - \rho^*(A)}{1+i_1} + \frac{\delta F\left(\rho^*(A)\right)(R - B/\Delta p)}{(1+i_1)(1+i_2)}}{1 - \frac{\delta f\left(\rho^*(A)\right)}{1+i_2}(R - B/\Delta p)}\right) d_A$$

The right-hand side of this inequality is maximized by setting $\rho*(A) = 0$, i.e. as low as possible, in which case $f(0) = \chi$ and $F(0) = 0$. Thus, in this way, the above inequality simplifies to

$$\frac{\chi}{(1-\delta)(1+i_2)} > \bar{A}(R - B/\Delta p)\chi \frac{\frac{r}{1+i_1}}{1-\frac{\delta\chi}{1+i_2}(R - B/\Delta p)}$$

Inspecting this result, we can see that if χ is large enough, namely larger than φ,

$$\varphi = \frac{\frac{r}{1+i_1}\chi(1-\delta)(1+i_2)\bar{A}(R - B/\Delta p)}{1-\frac{\delta\chi}{(1+i_2)}(R - B/\Delta p)}$$

then we have a sufficient condition for

$$\frac{\partial \int_{\bar{A}}^{I} d_l(A)dA}{\partial i_1} < 0$$

In order to prove that the aggregate change of LT debt of the unconstrained firms generalizes to individual firm behaviour, namely

$$\frac{\partial \int_{\bar{A}}^{I} d_l(A)dA}{\partial i_1} < 0 \ \Rightarrow \frac{\partial d_l(A)}{\partial i_1} < 0 \quad \forall A \in (\bar{A}, I]$$

we recall the assumption that unconstrained firms choose at least the minimum LT debt that allows them to equalize the highest LT debt share of the constrained firms, which is the LT debt share of the firm with endowment $\bar{A}$. This minimum component can be expressed as:

$$d_l^{min}(A) = \frac{(1+i_1)(I-A)}{\left[\frac{1-k}{k} + \frac{\delta F\left(\frac{\delta R}{1+i_2}\right)}{1+i_2}\right]}$$

where

$$k = \frac{R - B/\Delta p}{R - B/\Delta p + r - \frac{\delta R}{1+i_2}}$$

is the LT debt share of the firm with endowment $\bar{A}$. Additionally, we impose that for any unconstrained firm $d_l^{min}(A) < d_l(A) < (R - B/\Delta p)$, and that any change in the aggregate LT debt of unconstrained firms is distributed to all unconstrained firms as a change in their LT debt, and that this change, relative to the aggregate change, for any subset of unconstrained firms that has non-zero measure, is larger than zero.

Now, consider an infinitesimal increase in i_1: it will increase the minimum amount of LT debt needed to match the highest LT debt ratio of constrained firms.

Concretely:

$$\frac{\partial d_l^{min}(A)}{\partial i_1} = \frac{I-A}{\left[\frac{1-k}{k} + \frac{\delta F\left(\frac{\delta R}{1+i_2}\right)}{1+i_2}\right]}$$

However, as for all $A \in (\bar{A}, I]$ the choice of LT debt before the increase was strictly higher than the minimum LT debt due to our assumption, the infinitesimal change in the minimum LT debt will not make it surpass the previous amount. Instead, as the aggregate LT debt for the unconstrained firms must decrease, and this decrease is distributed to all unconstrained firms, we find that $\partial d_l(A)/\partial i_1 < 0$ for $A \in (\bar{A}, I]$. For an infinitesimal decrease in i_1, the aggregate increases while $d_l^{min}(A)$ decreases, thus in this case $d_l(A)$ must increase for all firms with $A \in (\bar{A}, I]$. This concludes the proof of statement 1.

Moving on, from

$$\frac{\partial \int_{\bar{A}}^{I} d_l(A)dA}{\partial i_1} = \frac{1}{F\left(\frac{\delta R}{1+i_2}\right)}\left[\frac{-\chi}{(1-\delta)(1+i_2)} - (R - B/\Delta p)\frac{\partial \int_0^{\bar{A}} F\left(\rho^*(A)\right) dA}{\partial i_1}\right]$$

we can see that if $\chi \to \infty$ then

$$\frac{\partial \int_{\bar{A}}^{I} d_l(A)dA}{\partial i_1} \to -\infty$$

Due to the assumptions that a change in the aggregate must be shared across firms and furthermore each share cannot be trivially small, by the same logic as above

$$\frac{\partial \int_{\bar{A}}^{I} d_l(A)dA}{\partial i_1} \to -\infty \Rightarrow \frac{\partial d_l(A)}{\partial i_1} < 0 \quad \forall A \in (\bar{A}, I]$$

For the constrained firms, $\rho^*(A)$ is pinned-down by the constraints, in which χ plays no role, as it does not affect the price. A change in χ thus has no effect on the choice of $\rho^*(A)$ for firms with $A < \bar{A}$. This concludes the proof of statement 2 of the proposition.

As shown in Lemma 2, constrained firms, those with $A < \bar{A}$, increase $\rho^*(A)$ in i_1. Their LT debt is determined by the constraint to be $d_l(A) = (R - B/\Delta p)$. This proves statement 3 of the proposition.

□

EXTENSIONS

EFFECT OF POLICY RATE CHANGE WITH RISK-NEUTRAL RATIONAL INVESTORS

To showcase the effect of our departure from a model with only rational and risk-neutral investors, we analyze a baseline specification, in which we assume that investors are risk neutral and rational: $\gamma = \alpha = 0$. In this case the investor demand function is horizontal, which means that they are willing to hold any amount of LT debt at price

$$P_l(A) = \frac{\delta F\left(\rho^*(A)\right)}{(1+i_1)(1+i_2)}$$

This is also the price at which a firm with endowment A would inelastically sell LT debt. Thus, the total amount of LT debt is not pinned down by market clearing. The requirements for constrained and unconstrained firms to exists are the same as derived above, and it is still true in this benchmark case that for constrained firms

$$\frac{\partial \rho^*(A)}{\partial i_1} < 0$$

However, because $\int_0^I P_l(A)\, d_l(A) dA$ is not pinned down by a downward sloping investor demand curve, we cannot say whether or how unconstrained firms adjust their maturity structure. Thus, we have that constrained firms take on less short-term debt when the monetary authority eases, while unconstrained firms have no incentive to change their maturity structure. This is counterfactual, as we see in the data that unconstrained firms should lengthen their maturity structure and do so more than constrained firms.

2. CAPITAL CONTROLS, CORPORATE DEBT AND REAL EFFECTS

Joint with Martha López Piñeros, José Luis Peydró and Paul Eduardo Soto

2.1 Introduction

Firms outside the U.S. have massively borrowed in dollars, especially in Emerging Markets (EM). Dollar credit to the non-bank sector outside the US amounted to 14% of global GDP in 2018, and EM debt accounts for roughly one third of the total value, with non-financial firms playing an important role in major EM (Aldasoro and Ehlers, 2018). Global banks –and local banks borrowing in dollars – have been key intermediaries for this increase in firms' foreign dollar funding (Bräuning and Ivashina, 2020, and forthcoming; IMF, 2019c). Cross-border loans, however, are especially fragile during financial downturns (De Haas and Van Horen, 2013; Giannetti and Laeven, 2012). Similarly, large capital inflows tend to precede credit booms, often followed by financial crises (Mendoza and Terrones, 2008; Reinhart and Reinhart, 2008; Jordà, Schularick, and Taylor, 2011; Gourinchas and Obstfeld, 2012). More generally, high corporate-leverage - especially if FX-financed - is a first-order risk for EM (Acharya et al., 2015; IMF, 2015; Alfaro et al., 2019; Bruno and Shin, 2019).

Capital controls after the last global financial crisis (GFC) have become increasingly popular among both policy-makers and academics, despite the historical costs associated to them (Johnson and Mitton, 2003; Rajan and Zingales, 2003), and the positive effects linked to financial liberalization (Henry, 2000a, 2000b). Even institutions such as the IMF have endorsed capital controls, though as a last-resort, temporary tool for managing credit booms led by large capital inflows, i.e. with a macroprudential type of role (IMF, 2012, 2018; Blanchard, 2013).[37] In the same spirit, a class of international finance-macro models rationalize capital controls as a Pigouvian tax to cut the negative externalities due to excessive foreign debt by firms (Bianchi, 2011; Brunnermeier and Sannikov, 2015; Jeanne and Korinek, 2010; Korinek, 2011).

We analyze the impact of capital controls on corporate debt and their real effects. For empirical identification: (i) we focus on the introduction (during a strong credit boom before the GFC) of a 40% *unremunerated* (at a time of very high interest rates) reserve requirement (URR) on foreign currency (FX) debt inflows in Colombia (capital controls (CC), Magud, Reinhart and Rogoff, 2011; Ostry et al., 2010); and (ii) we exploit matched administrative, proprietary datasets, including the supervisory credit registry and firm-level FX debt inflows and imports/exports (at quarterly frequency). The matched data allows us to study local and FX credit in

[37] EM have also supported capital controls, see e.g. https://www.ft.com/content/c27c016e-cf7e-11e8-a9f2-7574db66bcd5.

conjunction, and also the associated real effects (on firms' imports and exports) during the exogenous GFC, characterized by a world-level Great Trade Collapse (Bems, Johnson and Yi, 2013).

Briefly summarized, we find that capital controls reduce FX-debt inflows by 30%; with a further 10% cut for firms with one standard deviation higher ex-ante FX debt. Moreover, firms with ex-ante weaker relationships with local banks cannot substitute FX-debt with local debt (i.e. receive lower loan volume at higher loan rates, even controlling for firm fixed effects and other unobservables), thereby reducing firm-level total liabilities – and imports – immediately after the implementation of the policy. However, capital controls improve exports during the GFC (by 7.2% for an interquartile increase in exposure) by preemptively reducing firm-level total debt before the crisis, with stronger benefits for more ex-ante financially constrained firms (those with ex-ante tighter lending rates, maturity and collateral requirements). Importantly, benefits fully stem from reduction in corporate debt due to capital controls, not from endogenous changes in debt unrelated (orthogonal) to the policy. Results on both debt and trade are identical without controls or controlling for observables and a very large set of unobservables, thereby suggesting that selection is irrelevant for the results (Oster, 2017).

Our main contribution to the literature is to show how capital controls *benefit* the real economy via firms' capital structure – an *FX and local corporate debt channel mechanism* –; moreover, we exploit policy changes with administrative (local and FX) loan- and firm-level data for identification. Despite the increasing academic and policy attention on (prudential-type) capital controls and the large FX financing by firms, empirical evidence remains scarce, relying mostly on cross-country macro data (see, among others, Edwards, 2007; Forbes, Fratzscher and Straub, 2015; Zeev, 2017). Additionally, existing empirical literature on capital controls based on micro-data has focused on the *negative* effects, with either firm-level data (Johnson and Mitton, 2003; Desai, Foley and Hines, 2006; Forbes, 2007a, 2007b; Alfaro, Chari and Kanczuk, 2017) or loan-level data (Keller, 2019).[38] Interestingly, our results are different from the latter paper (using Peruvian policy and data), as Peru under capital controls allowed local banks to pass FX risk to firms, while Colombia did not. These different institutional details (and hence results) also show the limits of cross-country studies: specific regulations on controls are different, explaining why cross-country evidence is largely inconclusive (Magud, Reinhart and Rogoff, 2011). Moreover, by showing complementarities between FX debt and local (peso) credit supply, depending on the strength of local banking relationships, we also contribute to the large literature on lending relationships (Rajan, 1992; Petersen and Rajan, 1994; Bharath et al., 2007; Bebchuk and Goldstein, 2011; Bolton et al., 2016; Beck et al., 2018). The remainder of this Introduction is divided into two parts. First, we provide a detailed preview of the paper. Second, we discuss in detail the related literature and contrast it with our paper.

[38] Many papers highlight the positive effects of financial liberalization (see e.g. Henry 2000a, 2000b, and, from a long-run perspective, King and Levine, 2000, and Rajan and Zingales, 2003).

DETAILED PREVIEW OF THE PAPER

We investigate two main research questions. First, we ask whether, during the boom, the introduction of capital controls affect firms' FX and total debt and its potential consequences for the real economy. In detail, we analyze whether capital controls are effective in cutting FX-debt inflows, and also whether they are arbitraged away via domestic bank debt (and if so, the mechanism). Second, we analyze the potential positive real effects during the subsequent global financial crisis after the failure of Lehman Brothers in mid-September 2008 via a reduction of debt in the boom. That is, we analyze the effects of the capital controls from a prudential perspective during a boom and bust and investigate the debt channel as a potential mechanism.

Our work is primarily based on two administrative, confidential datasets. First, we have access to the National Credit Registry (CR), provided by the Colombian Financial Supervisory Authority, which collects detailed quarterly information at the loan-level for corporate loans, with information on loan volume, rates, collateralization, maturity, and currency. Differently from most credit registers around the world, we have loan rates which are important for isolating credit supply changes. Second, we exploit the Balance of Payments records on firm-level quarterly borrowing from foreign banks and in the form of trade credit and bond issuances, as well as firm-level quarterly imports and exports. Finally, we collect data on firms' and banks' (supervisory) balance sheet, with annual and quarterly frequency, respectively. All datasets are matched through firms' unique tax identifiers or through banking groups denomination codes.

For capital controls, we exploit the introduction of a 40% unremunerated URR on FX debt inflows by the Central Bank of Colombia in May of 2007 during a strong credit boom. At the time, local interest rates – as reflected by the overnight interbank rate – were as high as 8.40%. Hence, the new regulation resulted in high taxation of FX debt inflows as a large part of the inflows were in the central bank as unremunerated reserves. CC, which were borne by the ultimate borrower, were deposited for 6 months at the central bank without any remuneration; the deposit could be eventually withdrawn before this deadline, but against a heavy penalty fee. Importantly, FX-loans by local banks to firms (not only by foreign banks) were also taxed by the CC. The capital controls were lifted in early October 2008, amid signs of economic slowdown related to the unfolding of the GFC after Lehman's collapse.

We concentrate our analysis on 2,861 firms active in FX-debt markets before the URR.[39] Given both the introduction in May 2007 of the controls and the GFC after mid-September 2008, unless otherwise stated, we conduct our analysis of FX and total debt dynamics in 5-quarter symmetric windows around the policy introduction (i.e., the sample starts in 2006:Q1 - with 2007:Q2 labelled as the first year-quarter under capital controls - and ends in 2008:Q2 before the global crisis). Next, for analyzing the firm-level real effects during

[39] Conditional on issuing any foreign or domestic currency debt, FX-debt is on average 30% of total debt flows.

the global crisis, we expand our sample so to include the GFC. Our sample period is therefore 2006-2009, at quarterly level.

As capital controls are non-random, but rather induced by the credit boom that affect corporate debt and real activity, we exploit firm heterogeneity in difference-in-difference (DID) models, controlling for common (observed or unobserved) time-varying shocks. Moreover, as ex-ante different FX-debt levels or financial intermediaries for each firm are also not random, we perform the test for selection into the treatment developed by Oster (2017) (following the literature initiated by Altonji, Elder and Taber, 2005) in all the key steps of our analysis (used e.g. by Mian and Sufi, 2014, and Smith, 2016), i.e. in FX inflows, credit, and trade. In our setting, this exercise is very informative, as by saturating models with high-dimensional fixed effects (that control for time-varying unobservables) and by controlling for time-varying observables, there are very large changes in the R-squared relative to the baseline versions of our models to formally test for coefficient stability. Even under more demanding assumptions than those conventionally applied for performing the test, results suggest that self-selection is irrelevant for the effects observed due to the capital controls.[40]

Our main findings are as follows. We first establish that capital controls are effective in reducing FX-debt inflows (for ex-ante FX-active companies). Relative to the average FX-debt pre-policy exposure, capital controls reduce inflows by 30%. Moreover, the decline is stronger for ex-ante highly exposed firms: a 1 standard deviation (s.d.) increase over the mean implies an additional 10% cut. The reduction is effective for FX-loans granted by both global and local banks.[41]

The next step is understanding whether more affected firms substitute the forgone FX-debt with domestic (peso) loans from local banks.[42] It is important to stress that capital controls would apply on FX-debt irrespectively of the lender's nationality. Thus, we distinguish companies depending on whether they borrowed (pre-policy) in FX from local or foreign banks. We use this grouping to compare the relative performance in the domestic peso-lending market through credit register data. We find that after the implementation of the capital controls, companies *without* FX-lending relationships with local banks face a relative credit restriction of 13% vis-à-vis companies *with* ex-ante FX-relationships with local banks. The reduction in credit volume is accompanied by a relative interest rate jump of 71bp, suggesting that the credit changes across firms are (bank)

[40] At the time of the capital controls there was a change in traditional reserve requirements (based on bank deposits) on Colombian banks' funding. Given our granular data, we can isolate the effects of capital controls: (i) in loan-level regressions, where we exploit firm heterogeneity on ex-ante FX exposure, by applying bank*year-quarter fixed effects, hence fully controlling for any credit-supply variation connected to banks' idiosyncratic shocks, including the reserve policy ones; (ii) in firm-level models, by controlling for direct exposure to the reserve policy using banks' supervisory balance sheet data. Decisively, none of our results change (the estimated coefficient is identical) on the inclusion of such controls, or more generally, on other type of controls or fixed effects based on the results following Oster (2017)'s test.
[41] Results are robust (both for FX-debt flows from local and foreign banks) if we repeat the analysis over any symmetric window around the introduction of capital controls, including a 1-quarter exercise where we compare FX-debt flows in 2007:Q2 and in 2007:Q1.
[42] On the extensive margin, we find that the relative likelihood of issuing peso debt (against FX-debt) rises with capital controls and proportionally to pre-policy FX-debt exposure. Also, the share of FX-debt out of total debt issuance declines accordingly. Note that CC also tax FX lending by domestic banks.

supply-driven. In addition, the described relative credit supply cutback (expansion) is stronger among companies with larger ex-ante FX exposure to foreign (local) intermediaries, which predicts the extent of FX-debt reduction. Overall, our results are consistent with the ex-ante strength of local lending relationships. By borrowing in FX (in addition to pesos) from local banks, in fact, some companies become more transparent to the local banking system – as hard information on domestic FX-loans is recorded in the credit register – and build even stronger relationships with their own FX-lender, which will for instance receive additional soft information on the operations financed through FX-loans. Further corroborating the importance of local lending relationships, indeed, we find that the relative expansion in credit supply enjoyed by these firms is mostly operated by their local FX-lender, rather than by the remaining local banks from which they borrow only in pesos.

The loan-level findings are also confirmed when we aggregate to the firm-level. That is, firms with ex-ante weaker relationships with local banks cannot fully substitute FX-debt with domestic peso borrowing, so that capital controls constrain their total debt growth. Comparing annual balance sheet data for end of 2006 and end of 2007,[43] we find that these firms experience a relative average reduction of approximately 4.5% in total debt liabilities. With capital controls in place, more affected companies consistently reduce imports. In particular, an interquartile variation in exposure to capital controls (i.e. larger ex-ante FX-debt from foreign banks, i.e. weaker local banking relationships) implies a 4.4% fall in firm-level imports.

As the capital controls on FX inflows were introduced before the GFC (lifted in October 2008), we can analyze whether the pre-crisis reduction in total firm debt caused by the capital controls is beneficial during an exogenous external negative strong financial shock, by exploiting Lehman's failure. To this end, we additionally expand our sample from Lehman's failure to the end of 2009. Colombia did not have any sign of economic slowdown before the GFC at the end of 2008:Q3. Moreover, the GFC was characterized by a world trade collapse (exports and imports), and our matched administrative data have quarterly information for each firm on imports and exports.

Our results show that capital controls improve exports during the global financial crisis (and world trade collapse) by a preemptive reduction in firm-level debt before the crisis (and after the policy introduction). In particular, an inter-quartile increase in ex-ante exposure to the policy (whose related firms have a higher reduction in corporate debt pre-crisis) implies during the crisis higher exports growth by 7.2%.[44] The estimated coefficient remains the same without any control as compared to the case with all the controls despite the R-squared jumps by 84 p.p. Importantly, the results are fully stemming from reduction in firm debt due to the

[43] Results are virtually identical if we compare total liabilities in end of 2006 and in end of 2008. However, we prefer the end-of-2006-to-2007 regressions as the end of 2008 is characterized by the GFC.
[44] Exports are unaffected when the CC are enforced, i.e., before the crisis.

capital controls; differently, endogenous changes in corporate debt (between the CC policy introduction and the start of the GFC) unrelated (orthogonal) to capital controls do not affect trade during the crisis.

Estimated effects are stronger for ex-ante financially-constrained firms, in particular firms with ex-ante higher cost of loans, or with higher collateral requirements, or with greater reliance on short-term debt. Separating firms based on the median value of these proxies of financial constraints, we find that (an interquartile) more exposed firms to CC that ex-ante pledge high levels of collateral benefit with a 28% rise in exports. Similarly, for ex-ante high loan interest-rate and more short-term-debt firms, effects are stronger both statistically and economically and amount to a 10% and 13% increase, respectively, in correspondence of the interquartile jump in exposure to the policy.[45]

All in all, our results suggest that the real effects of capital controls are stronger during the crisis (benefits) than during the implementation (negative real effects), comparing the economic and statistical effects on exports and imports.[46] Note, however, that we are not making a welfare analysis, we are just reporting benefits (and some costs) of capital controls via the corporate debt channel.

CONTRIBUTION TO THE LITERATURE

Our main contribution to the literature is to show that capital controls also benefit the real economy, and the mechanism is via firms' capital structure – a *FX and local corporate debt channel mechanism*. In addition to the literature on international capital flows, firm FX debt and capital controls, we also contribute to the large literature on credit in general.

Despite the increasing attention on prudential capital controls by both academia and policy, empirical evidence remains scarce, relying mostly on cross-country macro data, with the typical identification problems.[47] These studies normally try to assess the effectiveness of controls in terms of reduced inflows and domestic credit (e.g. Edwards, 2007, and Forbes, Fratzscher and Straub, 2015). Moreover, Zeev (2017) documents that Emerging Economies employing capital controls on inflows experience milder output reactions to global financial shocks. On the other hand, existing studies on capital controls based on firm-level micro-data have mostly focused on the *negative* effects, studying stock returns, investment rates and financial constraints of

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[45] For comparison, the fall in imports after the implementation of the policy differs only among firms with high vs. low collateral requirements. The former reacts to an interquartile variation in exposure to the policy with an 11% reduction in imports, while for firms with low collateral requirements, the effect is insignificant and the coefficient is much smaller.

[46] In robustness, we collect quarterly data on employment (that are not available at firm-level) for 27 manufacturing industries (3-digit ISIC) and collapse at the industry*year-quarter level our firm-level information by taking weighted averages across the industry (there is not quarterly firm-level data on real effects except for exports and imports; and there is not investment either for firm or industry-level data at the quarter level). Repeating exercises that are identical in nature to those applied with firm data, we find that: i) binding exposure to capital controls implies a reduction of total liabilities; ii) similar to exports, capital controls have no impact during the implementation phase, but importantly they are beneficial during the global crisis, with an industry-level interquartile variation in exposure to policy boosting employment by 1.9%.

[47] For a detailed account of recent theoretical and empirical findings in the literature on capital controls, see Erten, Korinek and Ocampo (2019) and Rebucci and Ma (2019).

listed companies from Emerging Markets *during the phase of implementation* of the policy.[48] We contribute to this literature by showing the FX and domestic corporate debt channel as a mechanism associated with positive, prudential real-economy benefits of capital controls during an (exogenous) crisis, which are absent in the empirical literature,[49] as well as the analysis of capital controls on a large sample of non-listed companies (that tend to be more financially constrained).

Interestingly, our results are likewise very different from a recent paper on capital controls using credit register data. Keller (2019) documents an unintended consequence of Peruvian controls in 2011, namely an increase in domestic firms' debt dollarization and associated fragility during a subsequent sudden stop. Such negative effects are explained by the fact that capital controls inhibited Peruvian banks from investing local dollar deposits in global forward markets, so that they were consequently redirected towards non-exporting firms. Her results and ours are not directly comparable, because of the different institutional frameworks of the Colombian and Peruvian capital controls and other institutional settings. Colombian banks were at the time of CC (and still are) inhibited from raising dollar deposits from Colombian households and firms. Crucially, the Colombian controls applied to FX-debt granted by *both local and foreign* financial intermediaries.

Importantly, the joint reading of the two papers raises a warning against reliance on cross-country studies on capital controls and helps explaining why the related empirical evidence is largely inconclusive (Magud, Reinhart and Rogoff, 2011).[50] Such studies generally label policies with different legal and institutional arrangements as capital controls. However, the two credit papers (ours and Keller, 2019), each one with very different results, show that institutional details are of first-order importance for understanding how capital controls transmit to banks and non-financial borrowers.

We further contribute to (and build a bridge between) the literatures on capital inflows and bank credit by showing complementarities between FX debt and local banks' credit supply, depending on the strength of local banking relationships. First, we show the mechanism of the corporate debt channel for our results on capital controls, where both FX debt inflows to firms and local credit supply to firms matter. Second, we are not aware of other studies identifying a credit channel behind the transmission of capital controls to the real economy that levers firms' heterogeneity in terms of the strength of local lending relationships (Sharpe, 1990; Rajan, 1992; Hoshi, Kashyap and Scharfstein, 1991; Petersen and Rajan, 1994; Berger and Udell, 1995). In this respect, our study adds to the evidence on how relationship lending shields corporate credit during financial downturns (Bolton et al., 2016; Beck et al., 2018) and at the same time allows banks to more easily pick up the slack left over by other retrenching lenders (Bharath et al., 2007). Third, the previous result in conjunction with the finding that local credit supply depends on foreign FX-debt reduction (affected by CC) suggest

[48] See e.g. Johnson and Mitton (2003), Harrison, Love and McMillan (2004), Desai, Foley and Hines (2006), Forbes (2007a; 2007b) and Alfaro, Chari and Kanczuk (2017).

[49] Related to our findings, Tong and Wei (2010) report evidence of smaller stock price falls during the GFC for companies in less financially opened Emerging Economies, including Colombia.

[50] Ahnert et al. (2018) show that, after general FX macroprudential policies, banks on average pass FX-risk to firms.

strategic complementarities in cross-border and local lending (Bebchuk and Goldstein, 2011; and Vives, 2014). Both channels are absent in Keller (2019), who also uses credit register data.

We finally highlight two additional contributions stemming from our findings on real effects. First, our paper relates to a novel empirical literature that tries to quantify the real effects of macroprudential measures with micro-level data (e.g. Igan and Kang, 2011, and Jiménez et al., 2017). In the context of EM, as far as we are aware, the only study that looks directly at firms' activity in relation to macroprudential policy is Ayyagari, Beck and Martinez Peria (2018), who find in a cross-country setting that companies operating in countries with tighter macroprudential stance invest less on average. Relative to them, we focus on a specific policy – (macroprudential) capital controls – and analyze its effects during a boom and a bust. Second, by showing ramifications of capital controls on firm-level trade, our study adds to a relatively large body of papers on the impact of financial shocks on trade (e.g. Amiti and Wenstein, 2011; Chor and Manova, 2011). In this respect, the negative impact of capital controls on imports mirrors Alfaro and Hammel (2007)'s findings that financial liberalization spurs imports. Differently, our documented macroprudential benefits in terms of higher exports suggest that capital controls could have mitigated the Great Trade Collapse in EM.

The rest of the paper is organized as follows. Section 2 describes the policy and datasets. Section 3 presents the results of capital controls on FX debt inflows. Section 4 adds local bank credit supply. Section 5 presents the real effects during the boom and the bust. Section 6 concludes.

2.2 Institutional Settings and Data

2.2.1 Capital Controls on Capital Inflows in Colombia

The Colombian economy experienced a rapid expansion in the mid-2000s, with annual GDP growth above 4% in both 2004 and 2005. At least from early 2006, inflationary pressures further intensified due to a pronounced surge in domestic credit. The annual growth rate of commercial credit more than doubled throughout 2006, reaching a value of 22% at the end of the year from an initial point of less than 10% (Figure 1, Panel A). The Central Bank reacted by steadily increasing the interest rate, which jumped from 6% at the end of 2005 to 8% by early 2007, and further up to 10% in mid-2008. The tightening of monetary policy was accompanied by a reversal in the dynamics of net international portfolio flows, moving to strong capital inflows already by the third quarter of 2006 (Figure 1, Panel B).

To deal with the acceleration of domestic and foreign credit booms, the Central Bank resorted to capital controls on foreign inflows on May 7th, 2007, under the form of an Unremunerated Reserve Requirement (URR) on all new FX bank-loans granted to Colombian individuals and companies.[51] In practice, the URR works as follows: upon disbursement of the FX-credit to a Colombian firm, 40% of the nominal loan amount is deposited in an account at the Central Bank, without receiving any remuneration back. The deposit is always

[51] By May 23rd, the measure was extended to portfolio investments.

borne by the ultimate borrower of the debt (i.e. firms in our analysis) and can be withdrawn for free only after 6 months. At the time, local interest rates – as reflected by the overnight interbank rate – were as high as 8.40%. Hence, the new regulation resulted in high taxation of FX debt inflows.[52]

Importantly, firms would always pay the URR on FX-loans, independently of them being granted from local or foreign banks. Moreover, when local banks lend in FX, they finance such operations through FX-funding from abroad.[53] To avoid double taxation, local banks' FX-financing was thus exempted. Capital controls were enforced immediately upon announcement and eliminated by the 9th of October 2008, amid signs of economic slowdown related to the global unfolding of the financial crisis after Lehman Brothers' collapse.

Contemporaneously to the introduction of CC, the Central Bank also changed the regulation on traditional banks' reserve requirements, applying generally higher requirements on saving and checking deposits. Given our granular data, we can isolate the effects of capital controls from those of traditional banks' reserve requirements: (i) in loan-level regressions, where we exploit firm heterogeneity on ex-ante FX exposure, by applying bank*year-quarter fixed effects, hence fully controlling for any credit-supply variation connected to banks' idiosyncratic shocks, including the reserve policy ones; (ii) in firm-level models, by controlling for direct exposure to the reserve policy using banks' supervisory balance sheet data. Decisively, none of our results change based on the inclusion of such controls (or more generally due to other controls).

2.2.2 Data and Summary Statistics

Our work is primarily based on two administrative and confidential datasets observed during the period of interest 2006-2009. First, we have access to the National Credit Registry (CR) - provided by the Colombian Financial Supervisory Authority (Superintendencia Financiera de Colombia) –which collects detailed quarterly information at the loan-level on commercial debt outstanding. We aggregate information on size of the loan, collateralization and maturity at the firm-bank-currency level. The distinction across currencies is not available for loan interest rates, that are consequently available at the firm-bank level. Second, we observe Balance of Payments records on firm-level quarterly borrowing from foreign banks and in the form of trade credit (from foreign firms) and bond issuances. One key difference between these two datasets is that while CR-data refer to the firm-bank-currency stock of debt, we observe firm-level debt flows from abroad. We also obtain information on firm-level quarterly imports and exports.[54] Finally, we collect publicly available data on firms'

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[52] Earlier withdrawals were allowed but against the payment of a heavy penalty fee, decreasing in time and ranging from 9.4% of the deposit itself during the first month to 1.6% during the sixth and last month.

[53] Colombian banks, as banks from other countries which follow the Basel capital rules, basically fully hedge their FX-exposure. In fact, already before CC, banks could not have negative in-balance-sheet FX position, whereas the global net FX-position (comprehending off-balance-sheet assets and liabilities in FC) could not go below -5% of regulatory capital.

[54] Data on firms' employment are not accessible, hence we rely on figures for manufacturing industries that are released each trimester from the Colombian National Administrative Department of Statistics (DANE).

and banks' balance sheet, at annual and quarterly frequency, respectively. All datasets are matched through firms' unique tax identifiers or through banking groups denomination codes.

Our sample comprehends 2,861 firms active in FX-debt markets before the CC, excluding financial companies (ISIC codes 65 to 67) and utilities (ISIC codes 40 and 41). Unless otherwise stated, we conduct our analysis in 5-quarter symmetric windows around the policy introduction. That is, the sample starts in 2006:Q1 (with 2007:Q2 labelled as the first year-quarter under capital controls) and ends in 2008:Q2 before the crisis. We compute summary statistics over the pre-policy period 2006:Q1-2007:Q1 and report them in Table 1.

Panel A contains firm-level summary statistics. Regarding foreign inflows, the aggregate variable across local- and foreign-driven inflows, FX $\text{Inflows}_{f,yq}$, is given by the quarterly flow amount rescaled by total assets. This variable can take either positive or nil values, depending on whether FX-debt is issued or not, respectively. The presence of zeros and the rescaling by total assets produces small numbers in absolute value. This should not lead to underestimate the importance of FX-debt issuance for our companies, though. The variable Share-$\text{FX}_{f,yq}$ describes the fraction accounted for by FX debt flows out of total debt issuance. Conditional on issuing any foreign or domestic currency debt,[55] FX-debt represents on average around 30% of total debt flows. There are differences in the distribution of FX-debt inflows lent by local and foreign banks, FX-Local $\text{Inflows}_{f,yq}$ and FX-Foreign $\text{Inflows}_{f,yq}$. For both variables, we compute summary statistics over companies that have at least a positive entry during the pre-policy period. First, FX-lending relationships with local banks are more common (note the larger number of observations). In fact, 1,684 companies have FX-ties to local banks, whereas 402 companies borrow in FX from foreign banks and 775 firms enjoy FX-lending relationships with both local and global lenders. Second, foreign FX-debt flows are significantly larger. This reflects heterogeneity across firms borrowing in FX. Table 2 indeed indicates differences across companies in the two segments of the FX-debt market. Firms borrowing in FX from both local and foreign intermediaries are larger, with balance sheets around 1.5 and 0.8 times bigger than those of companies borrowing exclusively from local or foreign banks, respectively. The same ranking is also preserved along both imports and exports. One important remark is that all bank balance sheet characteristics are nearly identically distributed across the different groups of companies. This is a first reassurance that banks idiosyncratic characteristics do not interfere with the identification of the effects of capital controls based on the comparison between companies borrowing in FX.

A crucial variable in our analysis is the ex-ante exposure to FX-debt. Specifically, we aim to gauge a measure of pre-policy involvement in foreign currency borrowing. Since we do not have at our disposal the stock of foreign currency borrowing from abroad – in which case one might look at debt outstanding just at the onset of the policy, say in 2007:Q1 – we rely on a proxy given by the average issuance (rescaled by total assets) during the period from 2005:Q1 to 2007:Q1, the longest pre-policy period of observations for FX-inflows available to us. The related summary statistics for overall FX-debt exposure are those referring, in Table 1 and

[55] Note that this variable can be computed only for companies that issue at least one between peso and FX debt. For this reason, the number of observations for computation of statistics on *Share-*$FX_{f,yq}$ is lower.

2, to the variable $\text{Exposure}_{f,pre}$. Similar definitions apply to the exposures to FX-debt granted by local and foreign banks, respectively denoted by $\text{Exposure-Local}_{f,pre}$ and $\text{Exposure-Foreign}_{f,pre}$. Within subgroups of active companies, exposures contain heterogeneity. Across subgroups, firms with local FX-ties only are less reliant on FX-debt than the others, on average. Throughout the paper, we assess the robustness of our results to employing alternative measures of ex-ante exposure to FX-debt, which rescale inflows over total liabilities, or simply by taking logs, or consider their realization in 2007:Q1, or, finally, compute the average inflow over the period 2005:Q1-2005:Q4. We report their summary statistics in Table A1 of the Internet Appendix and they depict a substantially unmuted picture.

Firms total indebtedness is measured by its total liabilities, expressed in logs (of millions of Colombian pesos as of 2006:Q1, like other variables which are not rescaled by total assets) and denoted by the variable $\text{Liabilities}_{f,y}$, observed with annual frequency. Comparing the mean for total firm assets ($\text{Size}_{f,y-1}$) and liabilities, the latter account on average for 60% of a firm balance sheet.

The real effects of capital controls are analyzed over the period 2006-2009, so to study prudential benefits during the great financial crisis, exploiting quarterly data on imports and exports, expressed as well in logs and indicated by the variables $\text{Imports}_{f,yq}$ and $\text{Exports}_{f,yq}$, respectively. In exports (imports) regressions, we restrict our attention to those companies that during the period 2006-2009 export (import) in at least one year-quarter. For this reason, the number of observations drops, as not all companies in our sample engage in trade. Firms import more often than they export, which is reflected in fewer zeros. This also produces higher moments for imports than for exports.

Our analysis of the substitution of FX with local currency lending takes advantage of the credit registry, i.e. loan-level data. Panel B of Table 1 contains related summary statistics. The variable $\text{PesoLoan}_{f,b,yq}$ defines the log of the end-of-quarter firm-bank outstanding peso-denominated debt. The average peso-loan, expressed in end-of-2019 US dollars, is valued about $60,000.[56] The variable $\text{InterestRate}_{f,b,yq}$ represents the average interest rate applied over a company's debt balance with a given bank and is expressed in percentage points. The mean rate is 13.5%, reflecting the tight monetary policy stance of the Central Bank of Colombia over the period. Roughly 42% of the loans are collateralized and the average loan maturity is close to 4 years. Moreover, in 37% of the cases, a same bank grants not only peso credit, but also FX lending (as signaled by the variable $\text{FX-Lender}_{f,b,pre}$, a dummy with value 1 if a bank provides FX debt to a given firm before capital controls and 0 otherwise). Finally, note that firm-level variables are distributed differently in this sample for loan-level regressions. This reflects the fact that the number of firm-bank relationships is heterogeneously distributed across companies.

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[56] This figure is computed using the FRED CPI index for All Urban Consumers (https://fred.stlouisfed.org/series/CPIAUCSL) and the Peso-US$ exchange rate as of March 2006.

We report remaining summary statistics for macroeconomic controls and industry-level variables in Table A1 of the Internet Appendix.

2.3 Impact of Capital Controls on FX-Debt Inflows

We start our empirical analysis by looking at the influence of CC on FX-debt inflows. We study the behavior of the 2,861 ex-ante active companies in FX-debt markets during the period from 2006:Q1 to 2008:Q2. We intentionally exclude the third quarter of 2008 despite controls were effectively removed by early October of the same year. This is to separate the effects of capital controls from those of the GFC following the collapse of Lehman Brothers in mid-September of 2008, associated to high volatility of capital flows and to their retrenchment from EM towards Advanced Economies (Forbes and Warnock, 2012). All presented results nonetheless hold if we include 2008:Q3 in the regression sample (tables are available upon request).

First, we look at the unconditional impact of capital controls, by exploiting the following model:

$$\text{FX Inflows}_{f,yq} = \beta_1 \text{Post}_{yq} + \beta_2 \text{Macro}_{yq\text{-}1} + \beta_3 \text{Firm}_{f,yq-1} + \delta_q + \delta_f + \epsilon_{f,yq}$$

The dependent variable aggregates local-driven and foreign-driven FX-debt inflows;[57] later, we will consider both markets separately. The key parameter of interest is β_1, loading Post_{yq}, a dummy with value 1 starting from 2007:Q2, the quarter of introduction of the CC, and 0 before. Therefore, we analyze CC over 5-quarter windows before and after their introduction. We augment the model with quarter and firm fixed effects, δ_q and δ_f, controlling for quarter-specific shocks (i.e. seasonal effects) to FX-debt issuance and for time-invariant firm heterogeneity, respectively. In addition, we include a vector of time-varying macroeconomic controls, $\text{Macro}_{yq\text{-}1}$, comprehending: the lagged yearly variation of GDP and CPI index (i.e. yearly inflation); lagged values of the VIX and of the exchange rate, both expressed in logs, and of the monetary policy rate. We also augment the model with a battery of firm controls, including lagged values of firm size, ROA, imports, exports and firm-level weighted averages (across loans shares) of multiple bank balance sheet items – most notably, the share of assets accounted for by saving and checking deposits, that were differently affected from 2007:Q2 onwards. Standard errors are double-clustered at the firm and industry*year-quarter level.

We show results in columns (1) to (3) of Table 3. Column (3) displays the coefficients for the most robust version of the model which we just described. With capital controls in place, total FX-debt inflows are on average smaller by 0.004 (significant at 1% level). This coefficient is small in absolute terms, due to data on inflows being rescaled by total assets but still reflects a large effect of CC. In fact, comparing this number with firm-level summary statistics in Table 1, it equals 30% of the ex-ante mean FX-debt inflow (which, in turn, accounts on average for roughly 30% of total debt issuance). The effect is similar in columns (1) and (2), i.e.

[57] That is, the sum of FX bank loans, provided by local and foreign banks, bond issuance in FX and trade credit from foreign firms. Note that FX-bonds issuance and trade credit are tiny relatively to bank loans in our sample. For this reason, we normally refer to FX-bank loans and FX-inflows interchangeably.

in less saturated versions of the model. In Panel A of Table A2 of the Internet Appendix, we repeat the same analysis for different groups of companies, sorted according to whether they ex-ante borrowed in FX from: local banks (column 1); both local and foreign banks (column 2), or foreign banks only (column 3). The estimates for β_1 suggests that the unconditional reduction of debt inflows is similar across the groups of firms.

To check whether CC impact differently firms ex-ante more reliant on FX-debt, we next run the following regression:

$$\text{FX Inflows}_{f,yq}=\beta_1\text{Post}_t\text{*Exposure}_{f,pre}+\left(\beta_2+\beta_3\text{*Post}_t\right)\text{Firm}_{f,yq-1}+\delta_{i,yq}+\delta_f+\epsilon_{f,yq}$$

That is, we condition the effect of capital controls on the ex-ante FX-debt exposure, $\text{Exposure}_{f,pre}$. For easing comparison of the coefficients in columns 3 and 4, we de-mean such exposure variable. We now further include interacted industry and year-quarter fixed-effects, $\delta_{i,yq}$, controlling for time-varying industry-wide (ISIC 4-digit level) shocks. Firm controls are finally interacted with the Post_{yq} dummy, potentially allowing for different relations among firm characteristics and FX-debt intakes before and after the CC. Table 3, columns (4) to (10), shows the estimated coefficients, revealing that more exposed companies are more affected by the CC, as β_1 is negative and statistically significant.

About the economic significance of our estimates, considering the pooled estimates in column 7, for firms with FX-exposure 1 s.d. above the mean, there is an additional 0.0106 reduction in FX-debt inflows. Overall, this implies a total reduction close to 40% of their ex-ante FX-exposure, hence an additional 10% reduction relative to the average firm, based on the same metrics. In columns (8)-(10) of Table 3, we run separate regressions for different groups of companies, sorted depending on whether they ex-ante borrow in FX from local and/or foreign banks, and confirm results from pooled regressions.

We conclude this section with a list of robustness checks. First, differently FX-exposed companies may vary along dimensions that we do not control for through our set of controls and fixed effects. Among observable characteristics, for instance, FX-exposure positively correlates with firm size, which, in turn, may endogenously correlate with TFP growth. If this was a threat to our identification assumption – namely, the interaction between the Post_{yq} dummy and ex-ante FX-debt exposure being orthogonal to firm-specific unobserved time-varying shocks – we would observe instability of the coefficients of interest when adding controls and fixed effects. In this sense, we formally check the extent of self-selection along unobservables through the Oster (2017)'s test. Building on seminal work from Altonji, Elder and Taber (2005), she derives the proportional degree of selection into the treatment (relative to that inferred from the data) needed to nullify the estimated treatment effect, assuming a value $\tilde{R}^2$ for the hypothetical share of variance one would explain, were all the relevant residual heterogeneity controlled for. A "coefficient of proportionality" $\tilde{\delta}>1$ is interpreted as reassuring evidence, implying that further unobservable characteristics should correlate with treatment in a stronger manner than observables and unobservables captured by fixed effects. In Table A3 of the Internet

Appendix we provide the results of the test, both under the standard assumption that $\widetilde{R}^2 = \min\left\{1.3\widehat{R}^2; 1\right\} = 1.3\widehat{R}^2$, where $\widehat{R}^2 = 0.4615$ is the explained variability of column (7) of Table 3, and under the very restrictive assumption that $\widetilde{R}^2 = 1$. In both cases, the resulting degree of proportionality is strictly greater than 1.

Second, we analyze a relatively long 5-quarter window around the policy, so that results in Table 3 could in principle be driven by other events taking place either in 2006 or in late 2007 and/or early 2008. For this reason, we also consider all the shorter windows around the policy announcement. Estimates in Panel B of Table A2 display a persistently negative and statistically significant coefficient throughout all the different specifications.

Finally, we allow for different definitions of the exposure variables, including: values as of 2007:Q1; non-linear transformation of our averaged measure through log exposures; rescaling by total liabilities rather than by total assets; computation of average exposure over the period 2005Q1:2005Q4. All results go through (see Panel C of Table A2 in the Internet Appendix). All the discussed robustness exercises perform similarly when considering separate regressions for the different groups of companies. The related tables, not reported for brevity, are available on request.

Overall, this section has shown that CC drastically reduce the ability of all ex-ante FX-indebted companies to borrow in FX, and specially so for those that ex-ante heavily rely on this form of finance. Hence, for understanding whether CC affect firm activity, we need to quantify the extent to which corporates can substitute the forgone foreign currency debt with domestic peso lending.

2.4 Substitution of Foreign Debt with Domestic Bank Debt

We investigate such potential substitution in this section. First, we study substitution along the extensive margin and next over the intensive margin. As a summary of the results that we show in this section, all FX-active companies increase the frequency of issuance of peso debt after capital controls (relative to FX-debt), with a resulting descent in the share of FX-debt over total debt issuance. Nonetheless, on the intensive margin, following CC, firms with stronger FX-lending relationships with local banks enjoy much higher debt growth rates, relatively to other firms.

2.4.1 Impact of Capital Controls on Currency Composition of Corporate Debt Issuances

FX-debt intakes become much less frequent under capital controls. On the extensive margin, this can imply that ex-ante more FX-exposed companies issue domestic currency debt more frequently. We verify this hypothesis borrowing the identification strategy from Becker and Ivashina (2014). In detail, we retain firm*year-quarter pairs where either FX or peso-debt was issued, so to control for positive credit demand, while dropping those with no debt issuance or intakes of both types of financing, as they do not bring any information about the relative ability of companies to issue debt in different currencies.[58] The equation of interest takes the form:

$$DebtType_{f,yq}=\beta_1\ Post_t*Exposure_{f,pre}+\left(\beta_3+\beta_4*Post_t\right)Firm_{f,yq-1}+\delta_{i,yq}+\delta_f+\epsilon_{f,yq}$$

The dependent variable, $DebtType_{f,yq}$, is a dummy variable with value 1 if only debt issuance in peso is recorded and with value 0 in the opposite case where FX-debt is issued and peso debt is not. The saturation with fixed effects and controls mirrors the model for evaluating the impact of capital controls on debt inflows. In Table 4, columns (1) and (2) indicate that firms relatively more ex-ante reliant on FX-debt substitute more. Based on point estimates in column (2), a 1 interquartile jump in pre-determined exposure to FX-debt boosts the likelihood of issuing peso-debt by roughly 3.7%, corresponding to a 4.7% increase relative to the pre-policy average. Columns (3)-(5) report analogous figures for regressions run over separated samples for companies with local and/or foreign FX-ties.

This result points to a CC-induced drag on companies' debt-dollarization. We formally verify this hypothesis in columns (6)-(10), where we run a model with the share of FX-debt out of total debt issuance as dependent variable. The equation is otherwise identical to those analyzed so far, as long as right-hand side variables are concerned. Results indicate a decrease in the share of FX-debt over total debt issuance for more ex-ante FX-exposed companies. Results are again consistent across the three different groups. The presented findings differentiate the Colombian capital controls from the Peruvian case studied by Keller (2019) and, generally, from those FX-policies which put caps on banks' foreign currency funding and/or other investments different from lending, which tend to increase non-financial agents' usage of FX-loans (Ahnert et al., 2018).

2.4.2 Substitution with Peso Debt from Local Banks

For highly ex-ante FX-exposed firms, after capital controls the issuance of peso debt becomes more frequent and represents a larger share of total debt issuance. Nonetheless, it remains to understand whether the same firms also adjust on the intensive margin. To this end, we investigate loan-level data for loans denominated in pesos from the CR.

We contrast the post-CC dynamics in the domestic peso-credit market of the different groups of companies based on whether, before the policy, they borrowed in FX from local or foreign banks, or from both. A key observation is that borrowing in FX from domestic lenders grants a closer relationship with the local credit system. Locally issued FX-loans are in fact recorded in the CR, along with their entire credit history of repayments and defaults, whereas loans issued abroad are not. Moreover, the local FX-lender will also access additional soft information which is not recorded in the CR, therefore establishing an even tighter connection.

[58] Including firm*year-quarter pairs where both peso and FX-debt is issued, and coding the entry as peso issuance or FX-issuance based on the largest value among the two, does not alter results.

These differences are crucial for explaining our findings, that are presented in four subsections: first, we describe the empirical strategy for detecting relative changes in the volume and in the price of credit caused by capital controls; second, we report results from our baseline model; third, we perform a list of robustness exercises; fourth, we investigate a mechanism which explains our results.

EMPIRICAL MODEL

The companies are grouped into three categories according to the three following mutually exclusive 0/1 dummies. First, $Local_{f,pre}$ equals 1 for firms borrowing in FX before capital controls from local banks only. Second, $Foreign_{f,pre}$ has value 1 for firms ex-ante indebted in FX exclusively with foreign banks. Third, $Both_{f,pre}$ equals 1 for firms ex-ante borrowing in FX from both local and foreign banks.

Local represents the baseline group in the following regression:

$$Y_{f,b,yq}=\left(\beta_1\ Both_{f,pre}+\beta_2\ Foreign_{f,pre}\right)*Post_{yq}+\theta X_{f,b,yq}+\delta_{f,b}+\delta_{i,yq}+\delta_{b,yq}+\epsilon_{f,b,yq}$$

The dependent variable, $Y_{f,b,yq}$, is either the log of peso-loan provided by bank b to firm f, or the interest rate applied over it. β_1 and β_2 are the two parameters of interest, describing the post-capital controls dynamics of Both and Foreign firms in domestic credit markets, compared to Local. $X_{f,b,yq}$ is a vector of firm and loan-level controls. Firm controls include, on top of the usual variables applied in firm-level analysis, a dummy for whether a company defaulted in any loan over the past year. Loan Controls include a 0/1 collateralization dummy and the (log)-maturity of the loans. All controls are eventually fully interacted with the $Post_{yq}$ dummy. $\delta_{f,b}$ is a full set of interacted firm and bank fixed effects, controlling for firm-bank matching, whereas $\delta_{i,yq}$ are interacted industry and year-quarter fixed effects.

Peso lending may be impacted by the contemporaneous shock to banks' reserve requirements, rather than by (or in addition to) capital controls. In turn, this might generate a bias in our estimates if banks' sources of financing covary with companies' choice to participate in different FX-debt markets. Summary statistics in Table 2, however, tells us that this is not likely to be the case, as bank attributes are identically distributed across the different groups of companies. Still, there might be other unobserved banks' idiosyncratic shocks that differently affect the willingness of banks to extend credit to the various groups of companies before and after CC, for reasons that are unrelated to the CC themselves. Thanks to the granularity of our datasets, we directly tackle these concerns applying bank*year-quarter fixed-effects, $\delta_{b,yq}$, controlling for all time-varying (observed and unobserved) idiosyncratic bank shocks.

BASELINE RESULTS

Panel A of Table 5 contains the results from the estimation of the regression equation for loan quantity. The most robust specification is in column (5). Relative to firms borrowing ex-ante in FX exclusively from local banks, firms ex-ante indebted in FX only with foreign banks experience a credit reduction of about 13%.

Moreover, companies borrowing ex-ante in FX both from local and foreign banks suffer a halfway cut of 6.9%. Importantly, and confirming the exogeneity of participation into different FX-debt markets to banks heterogeneity, the coefficients magnitudes are virtually unaffected by including bank*year-quarter fixed effects, whose addition to the model also implies a tiny change in the R-squared; in other terms, the differences between the coefficients in columns (3) and (4) are not significant and bank time-varying heterogeneity explains a very small share of the relative changes in loan volume across companies (e.g. traditional RR do not affect the estimated coefficient nor add any statistical explanation).

Since we shut down Colombian banks' idiosyncratic shocks channel, we study the simultaneous loan interest rate dynamics across groups to understand whether changes in credit are driven by supply or demand channels. Panel B of Table 5 shows results for the model with loan interest rate as dependent variable. In column (5), which displays estimates for the most robust version of the model, the price of credit increases by 79bp (30bp) for firms ex-ante indebted in FX only (also) with foreign banks, relative to firms with ex-ante FX credit relationships exclusively with local banks. The joint reading of Table 5 and 6 reveals that the relative quantity and price of credit move in opposite directions after the implementation of capital controls: the suggested credit variations across groups of companies are therefore driven by supply factors, consistent with the strength of local lending relationships.

ROBUSTNESS

To start with, the consistency of our estimates depends on the validity of the parallel-trend assumption: absent capital controls, firms in different groups would have gone through parallel credit dynamics. In Figure 2, we depict the aggregate raw loan quantity across groups, normalizing it to 1 in 2007:Q1, the last quarter before the introduction of CC. Each group of companies experience positive credit growth before capital controls. After CC, however, only companies ex-ante indebted in FX exclusively with local banks remain on such increasing trend, with a decline for firms with no ex-ante FX credit from local banks and flat dynamics for companies borrowing in FX both locally and abroad. Similarly, in Figure 3, before the introduction of CC interest rate is on a rising path for all companies, with diverging dynamics following the implementation of CC (note that monetary rates were continuously increasing over 2006 to 2008, so rates go up always for all firms).

We also perform other robustness tests to ensure that CC drive results. We rely again on the Oster (2017)'s test to check whether self-selection into the treatment may potentially invalidate our findings. We run the exercise using two benchmarks for the hypothetical R-squared: first, the value associated to the inclusion of firm*year-quarter fixed effects, which would absorb all firm-specific time-varying shocks, i.e. the main candidates as potential omitted variables in our model; second, the usual upper bound at 1. The resulting proportionality coefficients are in Table A4 of the Internet Appendix and are both above 1 in quantity regressions. For price regressions, they are negative, suggesting that selection along unobservables reinforces the described patterns, if anything. In other terms, in this case, the correlation among residual unobservables and the treatment should have opposite sign than the correlation between observables (and unobservables controlled for by fixed effects) and the treatment itself.

On top of clustering standard errors at the firm-level in all CR regressions, as we exploit firm time-varying heterogeneity for our main coefficients of interest, we also collapse our observations in a firm-bank average pre/post dimension, following Bertrand, Duflo and Mullainathan (2004), and re-run our model. The main finding that companies which ex-ante borrow in FX only from foreign banks suffer a credit supply cut from local banks still applies (Table A5 of the Internet Appendix).

An additional sensitivity check regards the fact we observe interest rates at the firm-bank level, rather than at the firm-bank-currency level. For validating that results are driven by peso borrowing, we run the same regression on firm*bank*year-quarter triples with positive peso loans and no FX-debt. The results, available on request, confirm qualitatively and quantitatively those described for the larger sample.

MECHANISM

Building on the large literature on lending relationships, we investigate a mechanism for explaining our results that describes potential complementarities between domestic and external credit. Our test involves two steps. First, local FX-lending relationships are visible in the CR, and should therefore favor firms' ability to borrow in local markets proportionately to the overall exposure to the Colombian FX-debt market, proxied through $\text{Exposure-Local}_{f,pre}$. On the other hand, additional exposure to foreign banks, i.e. higher values of $\text{Exposure-Foreign}_{f,pre}$, might predict a marginal increase in the credit supply cut, as they make firms more opaque to the local banking system, generating complementarities between cross-border and domestic lending (Bebchuck and Goldstein, 2011; Vives, 2014).

Second, granting loans gives banks soft information about borrowers (which are not recorded in the credit registry). Hence, if FX-lending relationships are key for substitution, the relative credit expansion in favor of (ex-ante) FX-customers of local banks has to be operated more aggressively by their Colombian FX-lenders themselves.

We verify the first conjecture in column (6) of both panels of Table 5. Indeed, higher exposure to local (foreign) banks, i.e. *weaker* (*stronger*) relationships with the local banking system, grants greater (lower) levels of credit following capital controls, at relatively lower (higher) price. Quantitatively speaking, a 1 interquartile increase in ex-ante FX-exposure to local banks is associated with a 3.67% jump in credit and an interest rate descent of roughly 30bp. Conversely, a 1 interquartile increase in ex-ante FX-exposure to foreign banks is associated with a 2.77% decline in credit and a hike in interest rate of 12bp. Note that coefficients are remarkably stable in different and less saturated versions of the model and across different definitions of the variables for FX-exposures (see Table A6 and Table A7 of the Internet Appendix, respectively).

Finally, we confirm in Table 6 that the credit supply increase for companies borrowing in FX from local banks is driven by *their FX-lender(s)*. We perform the following exercise. Throughout the different regressions, we always maintain the group of companies with no ex-ante FX-debt from local banks. We compare the evolution of the price and quantity of their peso loans with those of peso loans granted to the other companies by the local FX-lender(s) (even columns) and by the rest of the banks (odd columns). Results indicate that the relative credit expansion (and contemporaneous price descent) experienced by companies borrowing in FX only from local banks is mostly driven by a change in supply of the local banks which provided FX-loans before CC.

Overall, the evidence in this subsection suggests a mechanism based on companies being penalized (favored) because of looser (stronger) relationships with the local credit system.

2.5 Real effects

In this section, we study whether capital controls impact the real economy through their influence on firm debt. In detail, we first check that capital controls impacted the growth of firms' total debt. Consistently with the evidence presented so far, we will confirm that this is the case for firms with weaker relationships with local banks, whose ex-ante exposure to FX-debt is ultimately constraining. Next, we exploit this heterogeneity to check real effects on trade at the firm-level.

Capital Controls were introduced in May of 2007 and removed in October of 2008. Interestingly, from our perspective, the lifting of the CC coincides with the eruption of the global financial crisis (GFC) beyond US borders due the collapse of Lehman Brothers. Note that the GFC was characterized by a world-level collapse in trade. Hence, exploiting our data on imports and exports, we can analyze not only the impact of the capital controls upon implementation, but also their prudential benefits, potentially associated to a preventive slowdown of debt growth just before a major financial crisis (a "corporate-debt channel").

2.5.1 Capital Controls and Reduced Growth of Total Liabilities

For understanding whether the CC have ramifications for the real economy, we first check that they affect the growth of firms' total debt. Companies with weak ex-ante credit relationships with local banks may be affected, as they suffer credit cutbacks from capital controls and are additionally penalized by their Colombian (peso) lenders. Note, however, that the negative credit supply shocks might have been compensated by an increase in other forms of financing such as trade credit provided by other Colombian firms.

We verify that this substitution mechanism is not sufficient to undo the documented debt reduction by analyzing the evolution of total firms' liabilities, whose information is unfortunately available only at annual frequency. This generates ambiguity for the definition of the timing of the CC, which were adopted in 2007:Q2 (and removed in 2008:Q3). We try to overcome it by taking a dual approach. First, we consider only end-of-2006 and end-of-2007 data, which is our preferred choice. By leaving out end-of-2008, in fact, we avoid confounding shocks associated to CC with those stemming from the GFC. Next, however, we also check that results hold in a different sample where we bring in observations for end-of-2008. This strategy allows to compare ex-ante and ex-post firm liabilities, though it is subject to the critique that end-of-2008 contains shocks due to the GFC. In practice, we show that irrespectively of the terminal year we consider in our sample, more ex-ante exposed companies to CC (through weak relationships with local banks and high FX-debt) experience a reduction in total liabilities.

We present results in Table 7. Here the $Post_{yq}$ dummy takes value 0 in 2006 and value 1 in subsequent years. In columns (1)-(6), the terminal year is 2007. First, we run a relative exercise across groups, and find that CC reduce total liabilities for companies with no ex-ante FX-lending relationships with local banks by 4.7% in the most robust version of the model in column (5), where we include all usual controls interacted with the post dummy and both firm and industry*year fixed effects. The reduction holds if we fix 2008 as the terminal year of the sample (column (7)). We also verify that the reduction in total liabilities is increasing along (constraining) exposure to the policy (through ex-ante higher foreign FX-debt inflows and weak lending relationships with local banks), consistently with the evidence from previous sections. Excluding 2008 from the analysis, the coefficients in column (6) reveals that an interquartile increase in pre-policy exposure to capital controls prompts an additional reduction in total liabilities of 1.05%. These figures nearly double in regressions where 2008 is the terminal year with CC in place.

Overall, the evidence presented in this subsection shows that capital controls ultimately cause a reduction in total debt growth for companies more ex-ante reliant of FX-debt and with weak ex-ante lending-relationships with local banks. We now verify whether such corporate-debt channel of capital controls has ramifications for the real activity.

2.5.2 Capital Controls and Trade during the Boom and the Bust

Figure 4 shows that aggregate-level Colombian trade grew at fast and stable annual rates, close to 20%, from 2006 to mid-2008. Nonetheless, posterior dynamics indicates that Colombian imports and exports were affected by the Great Trade Collapse associated to the GFC of 2008-2009 (Bems, Johnson and Yi, 2013). The timing of CC (introduced in the boom and removed just before the unfolding of the GFC), the global financial and trade shock, and the availability of administrative quarterly firm-level data on imports and exports allow us to ask whether CC smooth the contraction in trade associated to the GFC by preemptively reducing corporate debt.

In this section, we answer this question, presenting findings in favor of such hypotheses. First, we describe our empirical strategy. Second, we present the baseline results. Third, we provide evidence that results are driven by a corporate-debt channel mechanism through a direct test, based on the decomposition of variation in firm total debt over 2006-2007 (hence during CC and before the crisis) into a CC-related component and a more endogenous one, orthogonal to the introduction of CC. Fourth, we perform a list of robustness checks. Finally, we further investigate firms' heterogeneity in terms of financial constraints, providing additional evidence on the corporate debt channel.

EMPIRICAL MODEL

We extend our sample to include 2009, hence observations are now collected over the period 2006:Q1-2009:Q4. We exploit the following regression model at the firm*year-quarter level:

$$Y_{f,yq}=(\beta_1\ Post_{yq}+\beta_2 Crisis_{yq})Exposure\text{-}Foreign_{f,pre}+\left(\gamma_1+\gamma_2 Post_{yq}+\gamma_3 Crisis_{yq}\right)Firm_{f,yq-1}+\delta_{i,yq}+\delta_f+\epsilon_{f,yq}$$

The dependent variable is either imports or exports, defined in logs. Our aim is to measure how ex-ante binding exposure to the CC (*Exposure-Foreign*$_{f,pre}$) impacts firm-level trade both during the policy period (2007:Q2 to 2008:Q2) and during the crisis (2008:Q3 to 2009:Q4). To this scope, *Exposure-Foreign*$_{f,pre}$ is interacted with the $Post_{yq}$ and the $Crisis_{yq}$ dummies: the former has value 1 from 2007:Q2 onwards, the latter only starting from 2008:Q3.

The parameters of interest are β_1 and β_2, measuring the impact of exposure to capital controls on firm-level trade. In particular, β_1 describes the effect of capital controls during the phase of enforcement and relatively to the pre-CC period. β_2 estimates the effect of CC during the crisis, and relatively to the CC period. We include our standard set of firm controls, fully interacting them with the $Post_{yq}$ and $Crisis_{yq}$ dummies. In each regression, we will include the interacted ex-ante FX-debt exposure to local banks, not associated to reduced debt growth through capital controls and which should therefore not cause any real effect. Consistently with previous firm-level regressions, we saturate the model with firm and industry*year-quarter fixed effects, which is also the clustering-level of standard errors.

BASELINE RESULTS

Panel A of Table 8 contains the baseline results on firm-level trade. We focus our discussion primarily on columns (1) and (2). Firms with higher ex-ante FX-debt and strong FX-lending relationships with local banks do not adjust neither imports nor imports, both during the implementation of the CC and during the crisis, in line with our results that they could undo the external shocks due to CC through an increase of domestic credit supply.[59]

Higher exposure to capital controls (resulting from the combination of larger ex-ante FX-debt exposure and weak relationships with local banks), interestingly, delivers imports losses on impact (introduction of the policy), with a 1.38% (inter-quartile) increase associated to a marginal 4.4% fall. Note also that imports do not revert to pre-CC levels during the crisis. In contrast, exports are not affected upon implementation of the CC.

[59] Columns (1) and (2) exclude companies ex-ante borrowing in FX from both Colombian and foreign institutions as these confound the effects of our treatment variable. Such companies in fact experience a relatively milder credit cutback (see Table 5) and their total firm-level liabilities are not constrained (see Table 7). Hence, CC are not binding for debt growth and may not be associated to a corporate debt channel for the real effects of CC during the crisis.

However, during the global crisis, exposure to capital controls is beneficial, with an interquartile increase associated to a 7.2% jump in exports. In robustness checks below, we will show that both results on imports and exports are completely robust across different versions of the model, including one with no controls nor fixed effects, and, consistently with previous sections, we will verify formally this claim through the Oster (2017)'s test.

Before, however, one first interesting observation emerges from the regression for exports in column (3) where we include companies with ex-ante FX-ties both domestically and abroad: the benefits of ex-ante foreign FX-exposure during the crisis diminish. We interpret this finding as prima-facie evidence supporting our "debt channel" mechanism: as already mentioned, CC do not constrain the debt growth of the newly included companies, serving their "prudential" role imperfectly and bringing weaker benefits during the GFC.

Capital controls therefore come with costs and benefits. On one side, CC reduce imports; on the other side, exports are unaffected in the aftermath of the policy but grow relatively faster during the crisis. The magnitudes of the benefits during the bust outweigh those of the costs during the boom, though, as suggested by our discussion on the economic significance of the estimated coefficients. However, as we argue in the Introduction, our paper does not perform a welfare analysis: we just report benefits and (some) costs.

MECHANISM

We run a direct test for our mechanism, the corporate debt channel, based on the hypothesis that the pre-crisis reduction in total debt due to CC is beneficial and drives the relative increase in exports for exposed firms.

In particular, we verify that endogenous drops in total debt – i.e. cuts in total liabilities growth orthogonal to exposure to capital controls – are not associated to post-crisis differences in exports. Excluding endogenous effects of total liabilities reassures that our estimates reflect a corporate debt channel due to capital controls, rather than other spurious dynamics. The test involves two steps. First, we run a cross-sectional regression of yearly reduction in total liabilities (i.e., yearly growth rate with negative sign) as of end-of-2007 against exposure to capital controls and industry fixed effects. This model is similar, but not identical, to that we used in the estimates of Table 7 (column 6),[60] and produces comparable coefficients (with higher significance at 1% level). The predicted values from such regressions are denoted by $-\Delta_{1y}\text{Liabilities}_{f,2007}^{\text{predicted}}$: they represent the drop in total firm debt prompted by exposure to capital controls. The residuals from the same regression are indicated by $-\Delta_{1y}\text{Liabilities}_{f,2007}^{\text{residual}}$, and constitute the endogenous variation in total firm debt, orthogonal to CC by construction. In the second step, we replicate our model, though substituting exposure to CC with $-\Delta_{1y}\text{Liabilities}_{f,2007}^{\text{predicted}}$, and further including $-\Delta_{1y}\text{Liabilities}_{f,2007}^{\text{residual}}$ as an additional independent variable. Summary statistics for both variables are shown in Table A1 of the Internet Appendix.

[60] The only difference is the exclusion of firm controls, contributing marginally to the total variation in total liabilities.

Panel B of Table 8 shows the results. Perhaps not surprisingly, the coefficients suggest that the reduction in firm debt caused by capital controls is associated with benefits in terms of exports during the GFC. Importantly, the endogenous reduction in total liabilities (orthogonal to CC) does not affect exports, providing evidence in favor of the corporate debt mechanism.

ROBUSTNESS CHECKS

We perform a list of robustness checks, reported in Table A8 of the Internet Appendix.

First, in Panel A and B we report the model for exports and imports, respectively, under different and progressively saturated specifications. The described results persist from the most basic version of the model with neither controls nor fixed effects, to the most robust one in column (4), which mirrors Table 8.

We also formally test coefficient stability through the Oster's test. In particular, for exports (imports) regressions we run the test for the coefficient loading the interaction between the $Crisis_{yq}$ ($Post_{yq}$) dummy and the constraining exposure to CC, capturing the real benefits (costs) of the CC during the crisis (implementation of the policy). In both cases, we assume $\tilde{R}^2 = \min\left\{1.3\hat{R}^2; 1\right\} = 1$, where $\hat{R}^2$ is the R-squared from most saturated model (in column (4) of Panels A and B for exports and imports, respectively). We report the coefficients of proportionality in Panel C and they are both strictly above 1, with an especially high value of about 33 for exports regressions.

In Panel D, we check that results are robust to different definitions of the variables measuring ex-ante FX-debt exposures. Consistently with previous sections of the paper, we employ proxies which rescale inflows by total liabilities, or simply by taking logs, or consider realizations as of 2007:Q1, or, finally, compute the average inflow over the period 2005:Q1-2005:Q4. Results generally hold across alternative definitions.[61]

Additionally, we also collapse our observations as firm-level averages during the three periods of interest, following Bertrand, Duflo and Mullainathan (2004), and re-run our model. That is, for each firm, we compute the mean value of imports and exports, and of the left hand side variables as well, over the periods: 2006:Q1-2007:Q1 (pre); 2007:Q2-2008:Q2 (policy); 2008:Q3-2009:Q4 (crisis). In this framework, the dummy $Post_{yq}$ has value 0 during the pre-period and value 1 during the policy and crisis periods. Moreover, the dummy $Crisis_{yq}$ has value 1 during the crisis period and 0 otherwise. We report results in Panel E and they are both qualitatively and quantitatively similar to those from baseline regressions.

[61] Measuring exposures through the realization of locally or foreign-driven FX-inflows (rescaled by total assets) as of 2007:Q1 generates inconsistent results (relative to the baseline findings) for imports. However, for all other measures taking averages over longer periods, baseline results hold. Note that taking a single year-quarter realization of FX-inflows may be problematic, as flow variables do not add over time. As a result, a single entry may not appropriately reflect the FX-debt exposure of a given company.

In Panel F, we check the robustness of our results to different definitions of the crisis and of the policy periods. After all, the CC were lifted in early October 2008 and Lehman Brothers collapsed in mid-September of the same year. Therefore, at face value, we may label 2008:Q3 as a policy quarter (columns 1 and 2) or, alternatively, exclude it from the analysis (columns 3 and 4). In both cases, baseline findings are unaffected.

In Panel G, we exclude companies operating in sectors related to the extraction, production and processing of oil (broadly defined, these correspond to ISIC sectors 10, 11, 12, 13, 14, 23 and industries 2521, 2529 and 2924), which represents a high share of Colombian trade. One concern is that the finding are disproportionately linked to the behavior of oil-related companies, which might have experienced specific dynamics unrelated to CC while being at the same time exposed to them. Nonetheless, estimated coefficients reassure that oil companies are not driving our results.

In Panel H, we further include companies that do not borrow at all in FX, hence unaffected by the CC. Comparing their trade-performance with FX-indebted companies is therefore informative for isolating the effects of CC through the corporate debt channel. Indeed, results are both quantitatively and qualitatively unaffected, although the statistical significance of coefficients in the exports regressions goes down.

On a similar vein, in Panel I, we re-run the baseline regressions within the group of firms ex-ante indebted in FX with foreign lenders, i.e. the firms more constrained by capital controls. By doing so, we address further worries about firms' self-selection into different segments (local vs foreign) of the FX-debt markets, despite previous results on coefficients stability in Panels A, B and C suggest that self-selection does not drive results. In column 1, we report coefficients for the baseline version of the model for exports. Like in pooled regressions, exposure to controls has no impact during the phase of enforcement of CC and, at the same time, exerts benefits during the crisis. The usual interquartile increase in exposure to the policy boosts exports by 5.68% during the GFC. The coefficient is slightly smaller relative to the baseline version of the model, which is not surprising, given that the average company in the group is constrained by capital controls, so variation takes place just on an intensive margin. In column 2, we find again that benefits stem from variations in total debt caused by CC, rather than by endogenous changes in total debt orthogonal to the policy (which have zero effect). In columns 3 and 4, results for imports are comparable to those commented for pooled regressions.

Finally, in Table A9 of the Internet Appendix, we check that CC consistently impact other margins of firms' real activity. No other variables (such as investment or employment) are available at firm-level with quarterly frequency. Hence, we exploit industrial-level data on employment for 27 manufacturing industries (unfortunately, investment is also not available at industry level). We translate the approach followed so far at the firm-level at a less granular 3-digit industrial level.[62] For exposure variables, we collapse firm-level data

[62] The hypothesis that we test is whether capital controls, by reducing total debt growth, made companies more resilient to the crisis, with consequential effects at the industrial level. A key step, therefore, is to show that looser FX-ties to local banks constraint debt growth also at the industrial level. In the Internet Appendix, Figure A1, Panel A, suggests indeed that for the 27 industries that we match with firm-level data, the relation between exposure to capital controls and

by taking weighted industry-averages, with weights given by the size of a company's assets over total assets in the industry (as of end of 2006). We augment the model with the same firm controls[63] applied in previous sections and industry and year-quarter fixed effects. Estimates from the most robust version of the model in column (4) suggest that higher pre-policy exposure to CC increases employment during the crisis. In details, an inter-quantile variation in industrial pre-policy exposure to CC raises employment by 1.9% during the crisis (robust to other definitions of exposure to CC, i.e. proxies which rescale debt flows by total liabilities in column 5 or by taking logs in column 6). Also, confirming again firm-level evidence, CC do not affect employment after the implementation of the policy (i.e. before the GFC).

HETEROGENEITY

We test for further heterogenous effects of capital controls across companies. Under the new paradigms of capital controls, financially constrained companies benefit more from a preemptive reduction in debt growth, as they would otherwise find more difficult to refinance themselves during a negative financial shock, the downside being that upon implementation they might be affected in a stronger manner (see e.g. Korinek, 2011). Hence, we separate companies according to three proxies of ex-ante financial constraints derived from credit registry data: the interest rate paid on loans, the share of collateralized bank credit and the share of bank credit with short maturity (i.e., below or equal to 1 year). Note that companies with high interest rate are on average riskier. Similarly, high collateral requirements are normally applied to opaque and/or riskier companies, whereas companies relying extensively on short-term debt are more vulnerable to unexpected negative liquidity shocks. During an unexpected crisis, all these firms are likely to experience worse outcomes if their debt balance is relatively large. Hence, they are also supposed to benefit more from pre-crisis reduction in total indebtedness.

Before moving to the discussion of results, we describe how we build proxies of financing constraints. First, we run loan-level regressions of interest rate, collateralized-loan dummy and short-term-loan dummy against bank*industry*year-quarter fixed effects, over the period 2005:Q1-2007:Q1. The residuals reflect financial constraints which are due to firm-specific factors and "cleaned" from industry, lender-specific or common time-varying factors (and from all potential interactions among them). Then, in each year-quarter, we build a weighted firm-level average, with weights given by the loan share over total firm's banks credit. Finally, we compute the firm-level mean over the entire period.[64] We display results in Table 8, Panel C (Panel D) for

subsequent *reduction* in total liabilities between 2006 and 2007 is markedly positive. Note that such relation controls for industry and year fixed effects and is significant at the 1% level and is robust to the inclusion of firm controls. It implies a 5.8% reduction in total liabilities for a 1 interquartile increase in exposure to capital controls at industrial level. Furthermore, also at the industry level, like in firm-level analysis, ex-ante FX-exposure to local banks does *not* constrain total debt growth (Figure A1, Panel B).

[63] For time-varying firm controls, we take a similar approach and build time-varying weighted averages. All firm controls are interacted with the $Post_{yq}$ and $Crisis_{yq}$ dummies.

[64] Importantly, results presented below go through both if we build our measures based on the original loan rates, collateralization or short-term debt shares or on residuals derived from more saturated models (including for instance

exports (imports). Firms are split into highly- and lowly-constrained along the three margins taking the median value in the regression sample as a benchmark.[65] Since we lose few observations over the process, we make sure that baseline results for both exports and imports hold in the smaller samples we look at (see columns 1, 4 and 7 of Panels C and D of Table 8).

Regressions on exports suggest that the benefits of capital controls are concentrated among ex-ante more financially constrained companies. In detail, firms pledging ex-ante high levels of collateral benefit from an interquartile increase in exposure to capital controls with a 28% rise in exports (relative to a 7.2% average increase). Also, while benefits are not statistically significant among low interest-rate and low short-term-debt companies, they are both statistically and economically significant for constrained companies along both margins – and amount to 10% and 13%, respectively, in correspondence of an interquartile jump in exposure to the policy. Differently, the fall in imports during the implementation of the policy differs only among companies with high and low collateral requirements. In particular, the former react to an interquartile variation in exposure to CC with an 11% reduction in imports. For companies with low collateral requirements, the effect is not statistically significant and the coefficient is also much smaller. Overall, the evidence presented in this subsection suggests that the benefits of capital controls are larger among ex-ante more financially constrained companies, in line with the corporate debt channel documented in previous subsections.

2.6 Conclusions

In this paper we have provided a comprehensive empirical analysis of macroprudential capital controls. For empirical identification: (i) we focus on the introduction (during a strong credit boom and high interest rates) of a 40% *unremunerated* reserve requirement (URR) on foreign currency (FX) debt inflows in Colombia before the GFC, i.e. capital controls (CC); and (ii) we exploit matched administrative datasets, most importantly the credit registry and firm-level data on FX debt inflows and trade flows, all at quarterly frequency. Through these data, we study the dynamics of capital inflows and of the local credit cycle altogether and uncover a corporate debt channel through which capital controls impact the real economy.

Our robust results show that capital controls reduce FX-debt inflows (by 30%) and that the reduction is relatively stronger for firms with larger ex-ante FX borrowing (by further 10%). Crucially, not all the affected companies can substitute this credit cutback with lending in peso from domestic banks. In particular, firms with ex-ante relatively weaker relationships with Colombian banks suffer an additional restriction in credit supply and hence experience a slowdown in credit growth and total corporate debt. This corporate debt channel

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other loan characteristics). We also make sure that each of these methodologies work if we were to repeat them over the longer pre-crisis period 2005:Q1-2008:Q2. Related tables are available upon request.

[65] The residuals we use to build our measures of constraints represent the firms' specific differences relatively to the average values applied over loans granted in a given sector by a same bank in a specific year-quarter. Hence, an alternative reasonable choice is splitting companies based on whether their proxy is above or below zero. Firms with positive values are in fact more constrained than the average industry peer applying for a loan to a given bank over the pre-CC period. Indeed, results are robust to such specification and the tables are available upon request.

has real ramifications both during the phase of implementation of capital controls (the boom) and during the subsequent Great Financial Crisis (the bust). During the boom, firms more constrained by capital controls reduce imports. However, reduced debt growth in the boom grants a better performance during the bust, in the form of larger exports (by 7.2%), especially for financially constrained firms (between 28% and 10%). Effects during the crisis are fully stemming from a reduction in corporate debt associated to capital controls and not from endogenous debt change orthogonal to the policy (where the corporate debt changes are between the introduction of CC and the start of the GFC). Results on both debt and trade are identical without controls or controlling for observables and a very large set of unobservables, thereby suggesting that selection is irrelevant for the results (following e.g. Oster, 2017). For example, in the case of exports during the crisis, the estimated coefficient remains the same without any control as compared to the case with all the controls, despite that the R-squared jumps by 84 percentage points.

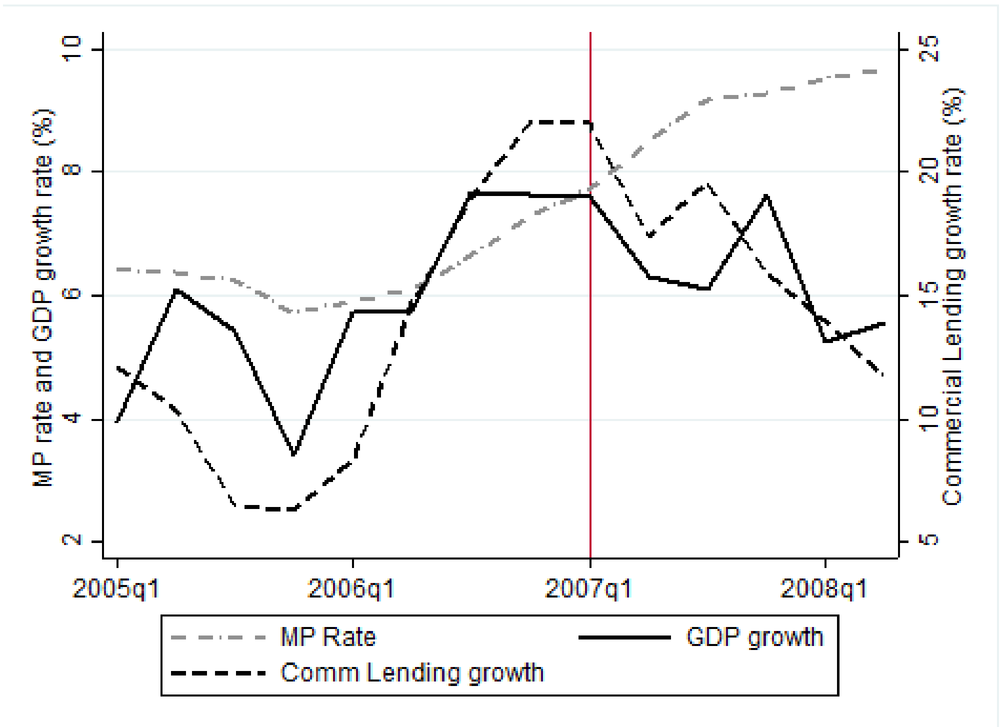
Our key contribution to the literature is to show benefits of capital controls for the real economy, starting from micro-level data (loan, firm and bank) and based on a corporate debt channel mechanism. This exploits the relative strength of firms' relationships with the local banking system as a channel for partly arbitraging the debt reduction from abroad due to the capital controls. Our results fill the gap between the increasing faith that both policy-makers and academics are arguing towards macroprudential capital controls and the inconclusive and problematic evidence based on time series and cross-country studies. Moreover, as we highlight twice in the Introduction, institutional details are crucial to understand the effects of capital controls (e.g. Keller (2019)'s results versus our results).

Finally, the literature has highlighted other channels through which capital controls may affect the real economy, including the strengthening of domestic monetary policy (Rey, 2015) and potential relations of complementarity/substitutability with other macroprudential measures (Korinek and Sandri, 2016). These questions are investigated in the next chapter of the thesis.

Figures

FIGURE 1: Macroeconomic Environment

Panel A: Credit Growth, Monetary Policy and Economic Growth



Panel B: Exchange Rate and Financial Flows

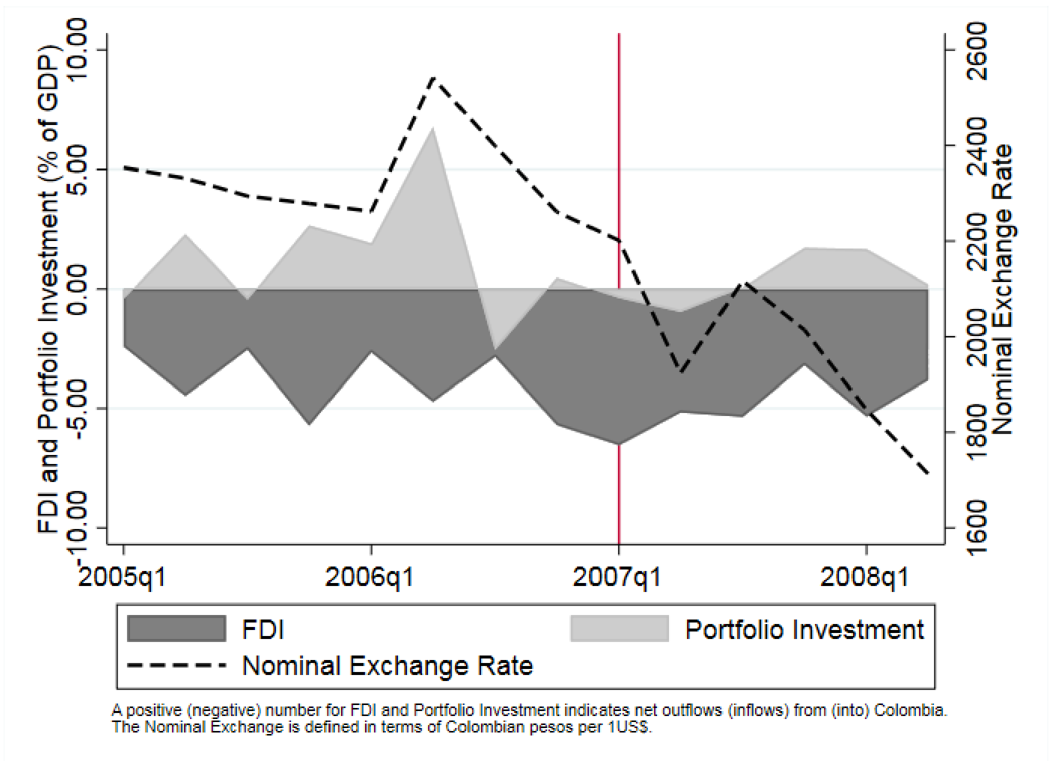


FIGURE 2: Aggregate Volume of Loans across groups of Companies

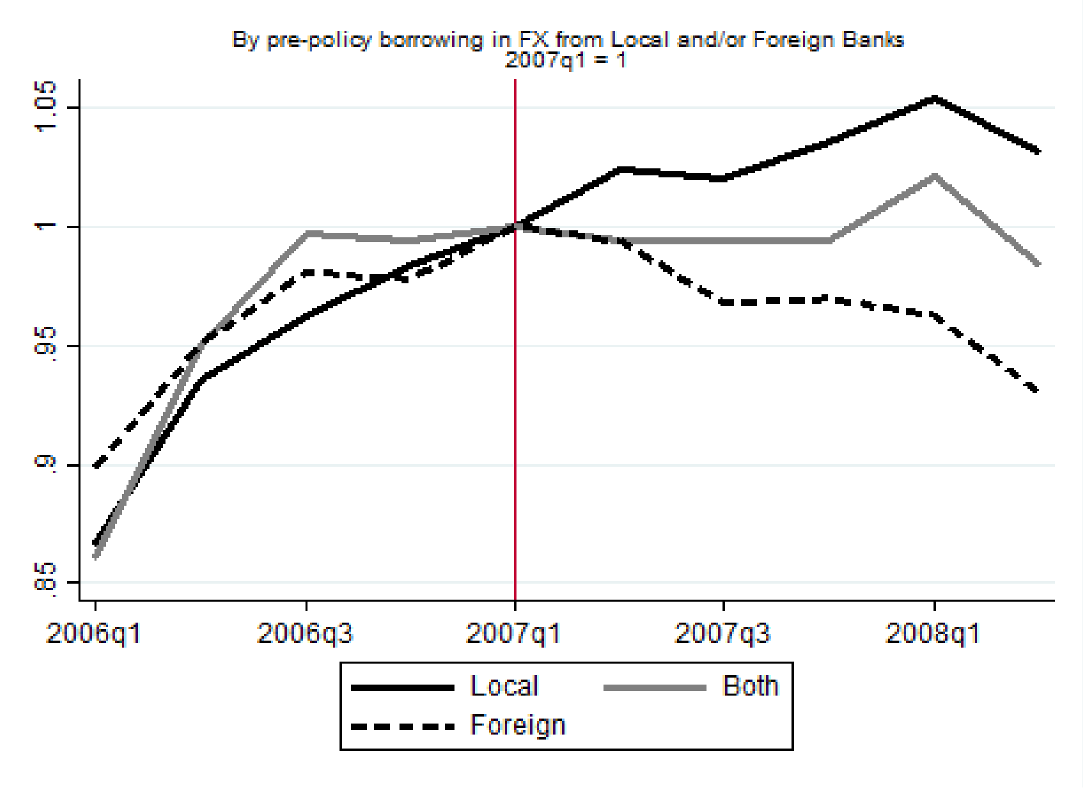


FIGURE 3: Average Loan Interest Rate across groups of Companies

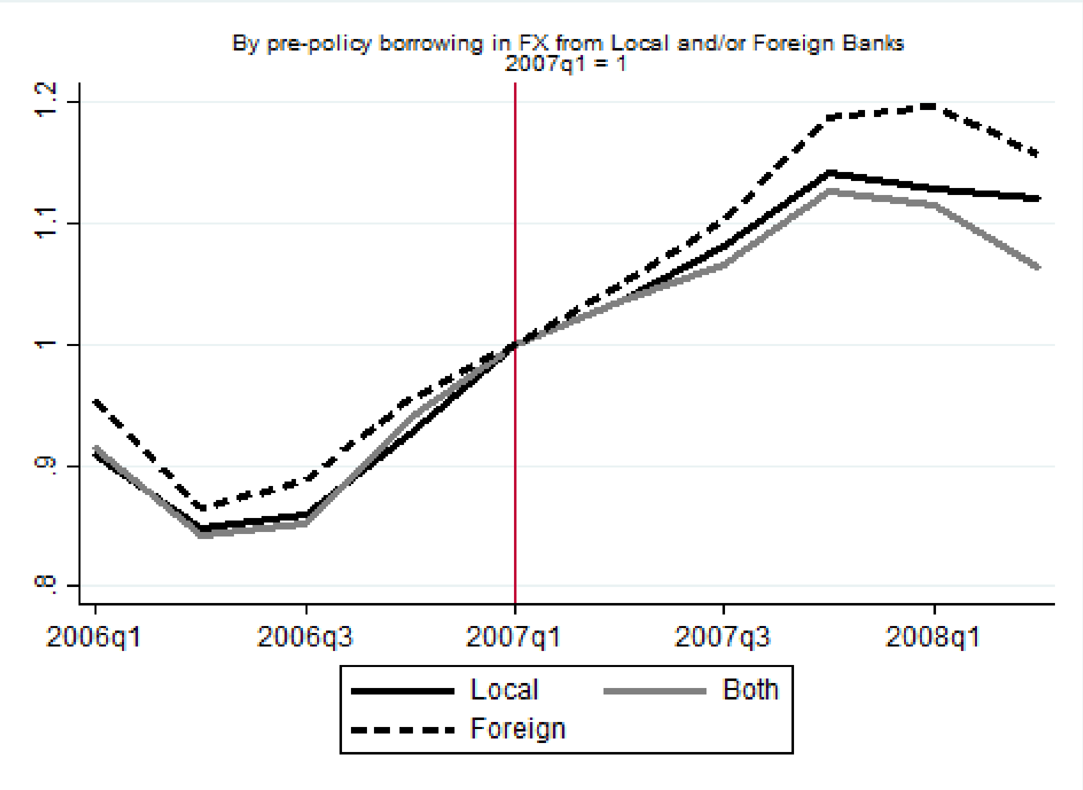
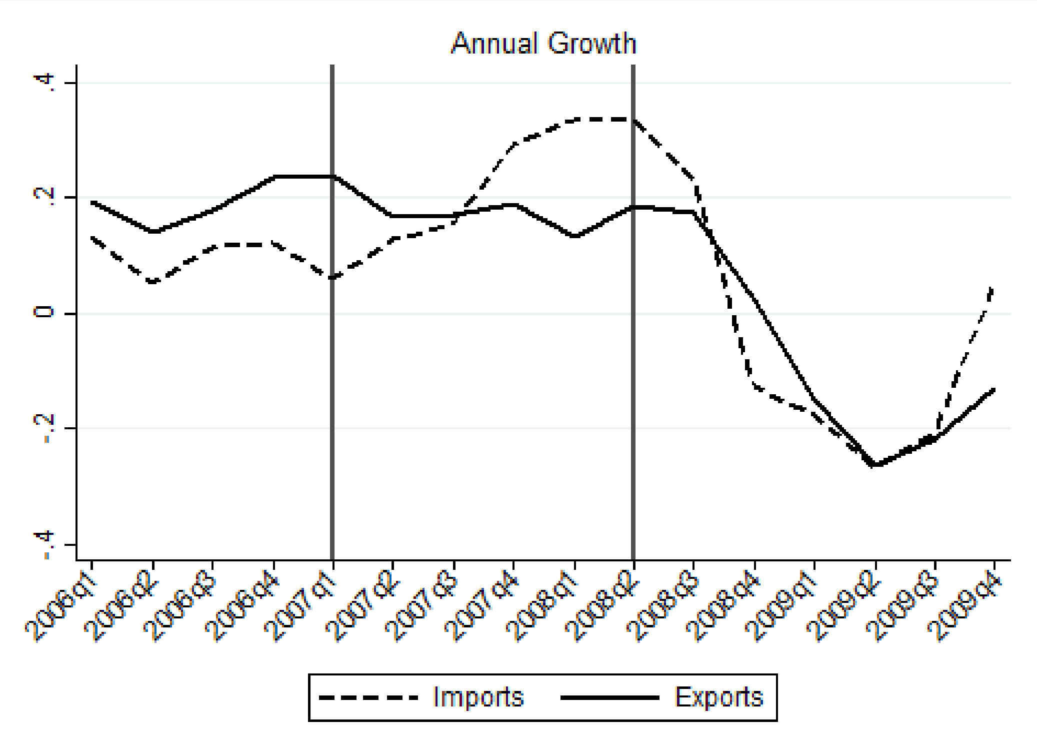


FIGURE 4: Country-level Imports and Exports



Tables

TABLE 1: Summary Statistics

PANEL A: Firm-level Analysis: 2006:Q1-2007:Q1

VARIABLES	Scale	N	Mean	P25	P50	P75	SD
FX Inflows$_{f,yq}$	Flow over Total Assets	14,125	0.0133	0	0	0.00660	0.0370
FX-Foreign Inflows$_{f,yq}$	Flow over Total Assets	5,751	0.0132	0	0	0.00351	0.0421
FX-Local Inflows$_{f,yq}$	Flow over Total Assets	12,176	0.00915	0	0	0.00216	0.0258
Share-FX$_{f,yq}$	$\in$ [0,1]	11,769	0.291	0	0.00371	0.631	0.397
DebtType$_{f,yq}$	0/1 Dummy	6,647	0.798	1	1	1	0.401
Exposure$_{f,pre}$	Flow over Total Assets	14,125	0.0132	0.000245	0.00374	0.0160	0.0231
Exposure-Local$_{f,pre}$	Flow over Total Assets	12,176	0.00853	0.000121	0.00152	0.00979	0.0151
Exposure-Foreign$_{f,pre}$	Flow over Total Assets	5,751	0.0143	0.00129	0.00506	0.0151	0.0257
Liabilities$_{f,y}$	Logs	14,125	8.374	7.198	8.318	9.512	1.673
ROA$_{f,y-1}$	Flow over Total Assets	14,125	0.0366	0.00931	0.0296	0.0627	0.0703
Size$_{f,y-1}$	Logs	14,125	8.848	7.678	8.802	9.952	1.621
Imports$_{f,yq}$	Logs	11,722	4.968	3.048	5.629	7.302	2.974
Exports$_{f,yq}$	Logs	7,938	4.074	0	4.512	7.021	3.362
BankCET1$_{f,yq-1}$	Stock over Total Assets	14,125	0.0397	0.0328	0.0388	0.0451	0.00865
BankROA$_{f,yq-1}$	Stock over Total Assets	14,125	0.0152	0.00960	0.0154	0.0197	0.00673
BankSize$_{f,yq-1}$	Logs	14,125	16.43	16.21	16.43	16.69	0.369
BankNPL$_{f,yq-1}$	Stock over Total Assets	14,125	0.0221	0.0197	0.0213	0.0235	0.00403
BankSaving$_{f,yq-1}$	Stock over Total Assets	14,125	0.334	0.303	0.331	0.361	0.0479
BankCheck$_{f,yq-1}$	Stock over Total Assets	14,125	0.146	0.125	0.140	0.165	0.0335
BankFX-Funds $_{f,yq-1}$	Stock /	14,125	0.0519	0.0392	0.0505	0.0638	0.0197
Default$_{f,yq}$	0/1 Dummy	14,125	0.0920	0	0	0	0.289
Relationships$_{f,yq}$	Discrete	14,125	3.816	2	4	5	1.996

TABLE 2: Summary Statistics – Firms Sorted by Pre-Policy Borrowing in FX from Local and/or Foreign banks

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	LOCAL (1684 companies)			BOTH (775 companies)			FOREIGN (402 companies)		
VARIABLES	Mean	P50	SD	mean	p50	SD	Mean	P50	SD
FX Inflows$_{f,yq}$	0.00826	0	0.0255	0.0229	0.00303	0.0468	0.0160	0	0.0505
FX-Foreign Inflows$_{f,yq}$	0	0	0	0.0118	0	0.0370	0.0160	0	0.0505
FX-Local Inflows$_{f,yq}$	0.00826	0	0.0255	0.0111	1.19e-06	0.0263	0	0	0
Share-FX$_{fx}$	0.235	0	0.378	0.396	0.261	0.403	0.301	0	0.422
Exposure$_{f,pre}$	0.00743	0.000736	0.0148	0.0237	0.0141	0.0291	0.0174	0.00541	0.0302
Exposure-Local$_{f,pre}$	0.00743	0.000736	0.0148	0.0109	0.00453	0.0155	0	0	0
Exposure-Foreign$_{f,pre}$	0	0	0	0.0128	0.00492	0.0228	0.0174	0.00541	0.0302
ROA$_{f,y-1}$	0.0437	0.0332	0.0714	0.0289	0.0251	0.0574	0.0211	0.0226	0.0832
Size$_{f,y-1}$	8.334	8.297	1.461	9.854	9.816	1.498	9.089	9.125	1.524
Imports$_{f,yq}$	2.850	0.774	3.136	6.028	6.745	2.773	4.716	5.485	3.136
Exports$_{f,yq}$	1.382	0	2.564	4.107	4.562	3.732	2.517	0	3.209
BankCET1$_{f,yq-1}$	0.0392	0.0381	0.00865	0.0403	0.0399	0.00799	0.0405	0.0396	0.00969
BankROA$_{f,yq-1}$	0.0154	0.0157	0.00672	0.0149	0.0148	0.00662	0.0150	0.0152	0.00693
BankSize$_{f,yq-1}$	16.46	16.46	0.363	16.40	16.39	0.336	16.36	16.40	0.436
BankNPL$_{f,yq-1}$	0.0220	0.0212	0.00402	0.0219	0.0213	0.00360	0.0227	0.0214	0.00478
BankSaving$_{f,yq-1}$	0.336	0.334	0.0484	0.329	0.326	0.0435	0.332	0.328	0.0526
BankCheck$_{f,yq-1}$	0.147	0.141	0.0337	0.144	0.139	0.0301	0.146	0.140	0.0384
BankFX-Funds$_{f,yq-1}$	0.0528	0.0510	0.0200	0.0520	0.0504	0.0180	0.0479	0.0478	0.0209
Default$_{f,yq}$	0.0774	0	0.267	0.111	0	0.314	0.117	0	0.321
Relationships$_{f,yq}$	3.631	3	1.889	4.635	4	2.127	3.012	3	1.614

LOCAL are companies that borrowed in FX only from local banks in the period from 2005:Q1 to 2007:Q1 and FOREIGN only from foreign ones. BOTH refers to the set of firms borrowing in FX from both local and foreign banks in the period from 2005:Q1 to 2007:Q1. Summary statistics are computed over the period: 2006:Q1-2007:Q1. FX Inflows$_{f,yq}$ represents total FX debt inflows, rescaled by total assets. FX-Foreign Inflows$_{f,yq}$ and FX-Local Inflows$_{f,yq}$ refer to FX-inflows intermediated by foreign and local banks, respectively, both rescaled by total assets. Exposure$_{f,pre}$ is the average of FX Inflows$_{f,yq}$ in the period from 2005:Q1 to 2007:Q1. Exposure-Local$_{f,pre}$ and Exposure-Foreign$_{f,pre}$ are the averages of FX-Local Inflows$_{f,yq}$ and FX-Foreign Inflows$_{f,yq}$ in the period from 2005:Q1 to 2007:Q1, respectively. Note: statistics on FX-debt flows intermediated by local and foreign intermediaries are computed over companies with at least one positive entry during the period 2005:Q1-2007:Q1. Share-FX$_{f,yq}$ is the share of FX-Debt flows out of total debt flows. Liabilities$_{f,yq}$ is the logarithm of firm. ROA$_{f,y-1}$ is previous year return on assets and Size$_{f,y-1}$ is the logarithm of total firm assets over the same period. Imports$_{f,yq-1}$ and Exports$_{f,yq-1}$ are the logarithm of (1 + firm imports) and (1 + firm exports), respectively. All variables with Bank prefix refer to firm-level weighted averages of local banks characteristics, where weights are loan share in total bank debt accounted for by a specific bank. BankCET1$_{f,yq-1}$ is bank common equity over total assets; BankROA$_{f,yq-1}$ is bank return on assets; BankSize$_{f,yq-1}$ is the logarithm of total bank assets; BankNPL$_{f,yq-1}$ is bank non-performing loans over total assets; BankSaving$_{f,yq-1}$ is bank saving deposits over total assets; BankChecking$_{f,yq-1}$ is bank checking deposits over total assets and BankFX-Funds$_{f,yq-1}$ is bank FX-liabilities rescaled by total assets. Default$_{f,yq}$ is a dummy with value 1 in case of firm default in at least one bank loan over previous year. Relationships$_{f,yq}$ is the number of local banks from which a company borrows.

PANEL B: Loan-Level Analysis (Regressions on Substitution of FX Debt with Peso Debt): 2006:Q1-2007:Q1

VARIABLES	Scale	N	Mean	P25	P50	P75	SD
Loan-level Variables							
Peso Loan$_{f,b,yq}$	Logs	50,527	5.145	3.836	5.349	6.758	2.233
Interest Rate$_{f,b,yq}$	%	50,527	13.57	9.400	13.42	18	7.142
Maturity$_{f,b,yq}$	Months	50,527	46.63	6	23.15	43.00	115.9
Collateral$_{f,b,yq}$	0/1 Dummy	50,527	0.422	0	0	1	0.494
FX-Lender$_{f,b,pre}$	0/1 Dummy	50,527	0.377	0	0	1	0.485
Firm-level Variables							
ROA$_{f,y-1}$	Flow over Total Assets	50,527	0.0337	0.00941	0.0278	0.0581	0.0636
Size$_{f,y-1}$	Logs	50,527	9.169	8.051	9.107	10.23	1.563
Imports$_{f,yq-1}$	Logs	50,527	4.381	0	5.262	7.227	3.359
Exports$_{f,yq-1}$	Logs	50,527	2.510	0	0	5.629	3.352
Default$_{f,yq}$	0/1 Dummy	50,527	0.111	0	0	0	0.314
Relationships$_{f,yq}$	Discrete	50,527	4.764	3	5	6	2.088
Exposure-Foreign$_{f,pre}$	Flow over Total Assets	50,527	0.0120	0.00124	0.00466	0.0120	0.0225
Exposure-Local$_{f,pre}$	Flow over Total Assets	50,527	0.00949	0.000156	0.00254	0.0114	0.0154

Summary statistics are computed over the period: 2006:Q1-2007:Q1. Firm-level Variables. FX Inflows$_{f,yq}$ represents total FX debt inflows, rescaled by total assets. FX-Foreign Inflows$_{f,yq}$ and FX-Local Inflows$_{f,yq}$ refer to FX-inflows intermediated by foreign and local banks, respectively, both rescaled by total assets. Exposure$_{f,pre}$ is the average of FX Inflows$_{f,yq}$ in the period from 2005:Q1 to 2007:Q1. Exposure-Local$_{f,pre}$ and Exposure-Foreign$_{f,pre}$ are the averages of FX-Local Inflows$_{f,yq}$ and FX-Foreign Inflows$_{f,yq}$ in the period from 2005:Q1 to 2007:Q1, respectively. Note: statistics on FX-debt flows intermediated by local and foreign intermediaries are computed over companies with at least one positive entry during the period 2005:Q1-2007:Q1. Share-FX$_{f,yq}$ is the share of FX-Debt flows out of total debt flows. Liabilities$_{f,yq}$ is the logarithm of firm. ROA$_{f,y-1}$ is previous year return on assets and Size$_{f,y-1}$ is the logarithm of total firm assets over the same period. Imports$_{f,yq-1}$ and Exports$_{f,yq-1}$ are the logarithm of (1 + firm imports) and (1 + firm exports), respectively. All variables with Bank prefix refer to firm-level weighted averages of local banks characteristics, where weights are loan share in total bank debt accounted for by a specific bank. BankCET1$_{f,yq-1}$ is bank common equity over total assets; BankROA$_{f,yq-1}$ is bank return on assets; BankSize$_{f,yq-1}$ is the logarithm of total bank assets; BankNPL$_{f,yq-1}$ is bank non-performing loans over total assets; BankSaving$_{f,yq-1}$ is bank saving deposits over total assets; BankChecking$_{f,yq-1}$ is bank checking deposits over total assets and BankFX-Funds$_{f,yq-1}$ is bank FX-liabilities rescaled by total assets. Default$_{f,yq}$ is a dummy with value 1 in case of firm default in at least one bank loan over previous year. Relationships$_{f,yq}$ is the number of local banks from which a company borrows. Loan-Level Variables. Peso Loan$_{f,b,yq}$ is defined as the logarithm of the loan in Pesos. Interest Rate$_{f,b,yq}$ is the interest rate paid on a given loan, defined in percentage points. Maturity$_{f,b,yq}$ is the maturity of the loan, in months. Collateral$_{f,b,yq}$ is a dummy variable with value 1 if a loan is collateralized and 0 otherwise. FX-Lender$_{f,b,pre}$ is a dummy variable with value 1 if bank b provides also FX debt (in addition to peso debt) to firm f between 2005:Q1 and 2007:Q1, and 0 otherwise.

TABLE 3: Impact of Capital Controls on FX-Debt Inflows

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
					FX $\text{Inflows}_{f,yq}$					
Post_{yq}	-0.003***	-0.003***	-0.004***	-0.005***	-	-	-	-	-	-
	(0.001)	(0.000)	(0.001)	(0.001)						
Post_{yq} *$\text{Exposure}_{f,pre}$				-0.429***	-0.429***	-0.459***	-0.461***	-0.401***	-0.377***	-0.533***
				(0.050)	(0.050)	(0.052)	(0.051)	(0.047)	(0.082)	(0.121)
N	28288	28288	28288	28288	28288	28288	28288	16394	7192	3317
R^2	0.0016	0.3903	0.3938	0.4149	0.4149	0.4167	0.4615	0.4748	0.5105	0.4954
Companies	All	All	All	All	All	All	All	Local	Both	Foreign
Firm FE	NO	YES	YES	YES	YES	YES	YES	YES	YES	YES
Quarter FE	NO	NO	YES	YES	-	-	-	-	-	-
Macro Controls	NO	NO	YES	YES	-	-	-	-	-	-
Firm Controls	NO	NO	YES	YES	YES	-	-	-	-	-
Bank Controls	NO	NO	YES	YES	YES	-	-	-	-	-
Time FE	NO	NO	NO	NO	YES	YES	-	-	-	-
Firm Controls*Post	NO	NO	NO	NO	NO	YES	YES	YES	YES	YES
Bank Controls*Post	NO	NO	NO	NO	NO	YES	YES	YES	YES	YES
Industry*Year-quarter FE	NO	NO	NO	NO	NO	NO	YES	YES	YES	YES

This table shows the effect of the introduction of capital controls on total FX debt inflows (rescaled by total assets), depending on pre-policy exposure to FX debt inflows. Post_{yq} is a dummy with value 1 from 2007:Q2 to 2008:Q2 and 0 from 2006:Q1 to 2007:Q1. $\text{Exposure}_{f,pre}$ is the average FX debt inflow (rescaled by total assets) over the period from 2005:Q1 to 2007:Q1. For easing comparisons between results in columns 4 and 5, we demean this variable. Macro Controls include lagged values of: GDP yearly growth rate; yearly inflation rate; log of VIX and of exchange rate and the lagged monetary policy rate. Firm Controls include $\text{ROA}_{f,y-1}$, $\text{Size}_{f,y-1}$, $\text{Imports}_{f,yq-1}$, $\text{Exports}_{f,yq-1}$. Bank Controls include: $\text{BankCET1}_{f,yq-1}$; $\text{BankROA}_{f,yq-1}$; $\text{BankSIZE}_{f,yq-1}$; $\text{BankNPL}_{f,yq-1}$; $\text{BankSaving}_{f,yq-1}$; $\text{BankChecking}_{f,yq-1}$ and $\text{BankFX-Funds}_{b,yq-1}$. The sign "-" denotes cases where a variable (or a group of variables or of fixed effects) is spanned out by other controls and/or fixed effects. Standard errors in parentheses are double-clustered at the firm and industry*year-quarter level. *** p<0.01, ** p<0.05, *p<0.1.

TABLE 4: Impact of Capital Controls on Currency Composition of Corporate Debt Issuances

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	$DebtType_{f,yq}$ (1=Peso; 0=FX)					$ShareFX_{f,yq}$				
$Post_{yq}$	0.019	-	-	-	-	-0.016	-	-	-	-
	(0.020)					(0.014)				
$Post_{yq}$*$Exposure_{f,pre}$	2.111^{***}	2.376^{***}	1.770^{**}	3.122^{***}	1.637^{**}	-1.691^{***}	-1.980^{***}	-2.134^{***}	-1.875^{***}	-1.627^{***}
	(0.346)	(0.385)	(0.850)	(0761)	(0.634)	(0.217)	(0.247)	(0.456)	(0.363)	(0.464)
N	13485	13485	8317	2384	1527	23278	23278	13181	6546	2237
R^2	0.3871	0.4846	0.4639	0.6022	0.6723	0.3594	0.4248	0.4187	0.4545	0.5970
Companies	All	All	Local	Both	Foreign	All	All	Local	Both	Foreign
Firm FE	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES
Quarter FE	YES	-	-	-	-	YES	-	-	-	-
Macro Controls	YES	-	-	-	-	YES	-	-	-	-
Firm Controls	YES	-	-	-	-	YES	-	-	-	-
Bank Controls	YES	-	-	-	-	YES	-	-	-	-
Year-quarter FE	NO	-	-	-	-	NO	-	-	-	-
Firm Controls*Post	NO	YES	YES	YES	YES	NO	YES	YES	YES	YES
Bank Controls*Post	NO	YES	YES	YES	YES	NO	YES	YES	YES	YES
Industry*Year-quarter FE	NO	YES	YES	YES	YES	NO	YES	YES	YES	YES

This table shows the effect of the introduction of capital controls on the relative frequency of peso vs FX debt issuance (columns 1 to 5) and on the share of FX debt out of total debt issuance (Columns 6 to 10), depending on pre-policy exposure to FX-debt market. Debt $Type_{f,yq}$ is a dummy with value 1 if a company issues peso-debt and value 0 if it issues: any FX-debt (columns 1, 2 and 4), local FX-debt (column 3) or foreign FX-debt (column 5). $Post_{yq}$ is a dummy with value 1 from 2007:Q2 to 2008:Q2 and 0 from 2006:Q1 to 2007:Q1. $Exposure_{f,pre}$ is the average FX-inflow (rescaled by total assets) over the period from 2005:Q1 to 2007:Q1. For easing comparisons between results in columns 1 and 2 and 6 and 7, we de-mean such variable. Macro Controls include lagged values of: GDP yearly growth rate; yearly inflation rate; log of VIX and of exchange rate and the lagged monetary policy rate. Firm Controls include $ROA_{f,y-1}$, $Size_{f,y-1}$, $Imports_{f,yq-1}$, $Exports_{f,yq-1}$. Bank Controls include: $BankCET1_{f,yq-1}$; $BankROA_{f,yq-1}$; $BankSIZE_{f,yq-1}$; $BankNPL_{f,yq-1}$; $BankSaving_{f,yq-1}$; $BankChecking_{f,yq-1}$ and $BankFX\text{-}Funds_{b,yq-1}$. Both Bank and Firm controls are fully interacted with the $Post_{yq}$ dummy. The sign "-" denotes cases where a variable (or a group of variables or of fixed effects) is spanned out by other controls and/or fixed effects. Standard errors in parentheses are double-clustered at the firm and industry*year-quarter level. *** p<0.01, ** p<0.05, *p<0.1.

TABLE 5: Substitution with Peso Debt from Local Banks

Panel A: Loan Volume

	(1)	(2)	(3)	(4)	(5)	(6)
				$PesoLoan_{f,b,yq}$		
$Post_{yq}$	0.259***	-0.087	-0.104	-	-	-
	(0.019)	(0.095)	(0.083)			
$Post_{yq}$* $Both_{f,pre}$	-0.093***	-0.077**	-0.062*	-0.064*	-0.069**	
	(0.035)	(0.038)	(0.035)	(0.035)	(0.035)	
$Post_{yq}$* $Foreign_{f,pre}$	-0.178***	-0.114**	-0.140***	-0.118***	-0.133***	
	(0.048)	(0.045)	(0.040)	(0.040)	(0.041)	
$Post_{yq}$*$Exposure\text{-}Foreign_{f,pre}$						-2.007*
						(1.134)
$Post_{yq}$*$Exposure\text{-}Local_{f,pre}$						3.793***
						(1.199)
N	102035	102035	102035	102035	102035	102035
R^2	0.044	0.258	0.789	0.791	0.802	0.802
Companies	All	All	All	All	All	All
Firm Controls*Post	NO	YES	YES	YES	YES	YES
Firm*Bank FE	NO	NO	YES	YES	YES	YES
Bank*Year-quarter FE	NO	NO	NO	YES	YES	YES
Industry*Year-quarter FE	NO	NO	NO	NO	YES	YES
Loan Controls*Post	NO	NO	NO	NO	YES	YES

Panel B: Loan Price

	(1)	(2)	(3)	(4)	(5)	(6)
			$InterestRate_{f,b,yq}$			
$Post_{yq}$	2.943***	6.683***	6.564***	-	-	-
	(0.073)	(0.373)	(0.373)			
$Post_{yq}$*$Both_{f,pre}$	-0.559***	0.365***	0.377***	0.327***	0.305**	
	(0.121)	(0.133)	(0.134)	(0.124)	(0.128)	
$Post_{yq}$*$Foreign_{f,pre}$	0.272	0.429**	0.358*	0.707***	0.786***	
	(0.190)	(0.192)	(0.186)	(0.170)	(0.170)	
$Post_{yq}$*Exposure-$Foreign_{f,pre}$						9.103**
						(3.552)
$Post_{yq}$*Exposure-$Local_{f,pre}$						-30.710***
						(4.135)
N	102035	102035	102035	102035	102035	102035
R^2	0.052	0.094	0.536	0.609	0.624	0.625
Companies	All	All	All	All	All	All
Firm Controls*Post	NO	YES	YES	YES	YES	YES
Firm*Bank FE	NO	NO	YES	YES	YES	YES
Bank*Year-quarter FE	NO	NO	NO	YES	YES	YES
Industry*Year-quarter FE	NO	NO	NO	NO	YES	YES
Loan Controls*Post	NO	NO	NO	NO	YES	YES

This table shows the effect of capital controls on the quantity and price of commercial (peso) credit granted from Colombian banks. In Panel A, the dependent variable is defined as the logarithm of the loan in pesos granted from bank b to firm f in year-quarter yq. In panel B, the dependent variable is the interest rate (in %) applied over the same loans. In columns (1) to (5), the baseline category is given by companies borrowing in FX before 2007:Q2 from local banks only. $Foreign_{f,pre}$ is a dummy with value 1 if a company borrowed in FX only from foreign intermediaries before 2007:Q2 and 0 otherwise. $Both_{f,pre}$ refers to companies resorting to both local and foreign intermediaries for peso credit before 2007:Q2. $Post_{yq}$ is a dummy with value 1 from 2007:Q2 to 2008:Q2 and 0 from 2006:Q1 to 2007:Q1. In column (6), Exposure-$Foreign_{f,pre}$ and Exposure-$Local_{f,pre}$ are the average of FX-Foreign Inflows $_{f,yq}$ and of FX-Local Inflows $_{f,yq}$ in the period from 2005:Q1 to 2007:Q1, respectively. Firm Controls include $ROA_{f,y-1}$, $Size_{f,y-1}$, $Imports_{f,yq-1}$, $Exports_{f,yq-1}$, $Default_{f,yq}$ and $Relationships_{f,yq}$. Loan Controls include: $Maturity_{f,b,yq}$ and $Collateral_{f,b,yq}$. Both Firm and Loan controls are fully interacted with the $Post_{yq}$ dummy. The sign "-" denotes cases where a variable (or a group of variables or of fixed effects) is spanned out by other controls and/or fixed effects. Standard errors in parentheses are clustered at the firm level. *** p<0.01, ** p<0.05, *p<0.1.

TABLE 6: Substitution with Peso Debt from Local Banks: Role of Ex-Ante FX Lending Relationships

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	$PesoLoan_{f,b,yq}$				$InterestRate_{f,b,yq}$			
	FX-Lender		FX-Lender		FX-Lender		FX-Lender	
	No	Yes	No	Yes	No	Yes	No	Yes
$Post_{yq}$*$Both_{f,pre}$	-0.041	-0.101*			0.206	0.362*		
	(0.039)	(0.055)			(0.150)	(0.216)		
$Post_{yq}$*$Foreign_{f,pre}$	-0.066	-0.238***			0.360**	0.851***		
	(0.043)	(0.051)			(0.181)	(0.221)		
$Post_{yq}$*Exposure-$Foreign_{f,pre}$			-2.802**	-0.110			11.316***	9.586**
			(1.348)	(1.133)			(4.247)	(4.415)
$Post_{yq}$*Exposure-$Local_{f,pre}$			-0.697	4.432***			-22.554***	-39.828***
			(1.794)	(1.482)			(5.367)	(5.558)
N	64443	48895	64443	48895	64443	48895	64443	48895
R^2	0.841	0.779	0.841	0.779	0.667	0.614	0.668	0.615
Firm Controls*Post	YES	YES	YES	YES	YES	YES	YES	YES
Firm*Bank FE	YES	YES	YES	YES	YES	YES	YES	YES
Bank* Year-Quarter FE	YES	YES	YES	YES	YES	YES	YES	YES
Industry*Year-Quarter FE	YES	YES	YES	YES	YES	YES	YES	YES
Loan Controls*Post	YES	YES	YES	YES	YES	YES	YES	YES

This table shows the importance of FX lending relationships with local banks for substituting FX-debt with peso-debt during capital controls. The samples vary across columns. We always keep all companies borrowing in FX exclusively from foreign banks. For the other companies: in even columns (FX-Lender: "Yes") we retain peso-credit relationships with Colombian banks that do provide FX-debt between 2005:Q1-2007:Q1; in odd columns (FX-Lender: "No"), with Colombian banks that do not provide FX-debt between 2005:Q1-2007:Q1. In columns (1) to (4), the dependent variable is defined as the logarithm of the loan in pesos granted from bank *b* to firm *f* in year-quarter *yq*. In columns (5) to (8), the dependent variable is the interest rate (in pp) applied over the same loans. In columns (1)-(2) and (5)-(6), the baseline category is companies borrowing in FX before 2007:Q2 from local banks only. $Foreign_{f,pre}$ is a dummy with value 1 if a company borrowed in FX only from foreign intermediaries before 2007:Q2 and 0 otherwise. $Both_{f,pre}$ refers to companies resorting to both local and foreign intermediaries for peso credit before 2007:Q2. $Post_{yq}$ is a dummy with value 1 from 2007:Q2 to 2008:Q2 and 0 from 2006:Q1 to 2007:Q1. In columns (3)-(4) and (7)-(8), *Exposure-*$Foreign_{f,pre}$ and *Exposure-*$Local_{f,pre}$ are the average of *FX-Foreign* $Inflows_{f,yq}$ and of *FX-Local* $Inflows_{f,yq}$ in the period from 2005:Q1 to 2007:Q1, respectively. Firm Controls include $ROA_{f,y-1}$, $Size_{f,y-1}$, $Imports_{f,yq-1}$, $Exports_{f,yq-1}$, $Default_{f,yq}$ and $Relationships_{f,yq}$. Loan Controls include: $Maturity_{f,b,yq}$ and $Collateral_{f,b,yq}$. Both Firm and Loan controls are fully interacted with the $Post_{yq}$ dummy. The sign "-" denotes cases where a variable (or a group of variables or of fixed effects) is spanned out by other controls and/or fixed effects. Standard errors in parentheses are clustered at the firm level. *** p<0.01, ** p<0.05, *p<0.1.

TABLE 7: The Impact of Capital Controls on Total Liabilities

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
				Ln(Total Liabilities)$_{f,y}$				
$Post_{yq}$	0.148	0.148^{***}	-0.283	-	-	-	-	-
	(0.150)	(0.013)	(0.656)					
$Post_{yq}$*$Both_{f,pre}$	-0.031	-0.031^{*}	0.003	0.003	0.005		-0.001	
	(0.227)	(0.016)	(0.017)	(0.017)	(0.016)		(0.020)	
$Post_{yq}$*$Foreign_{f,pre}$	-0.073	-0.073^{***}	-0.052^{**}	-0.052^{**}	-0.047^{**}		-0.043^{*}	
	(0.152)	(0.021)	(0.020)	(0.020)	(0.018)		(0.023)	
$Post_{yq}$*Exposure-$Local_{f,pre}$						0.692		0.254
						(0.496)		(0.660)
$Post_{yq}$*Exposure-$Foreign_{f,pre}$						-0.768^{*}		-1.418^{*}
						(0.428)		(0.851)
N	5632	5632	5632	5632	5632	5632	5616	5616
R^2	0.1705	0.9873	0.9878	0.9878	0.9881	0.9881	0.9767	0.9767
Firm FE	NO	YES	YES	YES	YES	YES	YES	YES
Firm Controls*Post	NO	NO	YES	YES	YES	YES	YES	YES
Bank Controls*Post	NO	NO	YES	YES	YES	YES	YES	YES
Year FE	NO	NO	NO	YES	-	-	-	-
Industry*Year FE	NO	NO	NO	NO	YES	YES	YES	YES
Terminal year	2007	2007	2007	2007	2007	2007	2008	2008

This table shows the effect of capital controls on total liabilities, depending on pre-policy firms activity in local/foreign FX-debt markets and on exposure to FX-debt markets. The dependent variable is defined as the logarithm of total liabilities of firm f in year y. In columns (1)-(6), observations are from 2006 and 2007. In columns (7)-(8), the sample includes observations for 2006 and 2008. In columns (1)-(5) and (7), the baseline category is companies borrowing in FX before 2007:Q2 from local banks only. $Foreign_{f,pre}$ is a dummy with value 1 if a company borrowed in FX only from foreign banks before 2007:Q2 and 0 otherwise. $Both_{f,pre}$ refers to companies resorting to both local and foreign banks for FX-credit before 2007:Q2. In columns (6) and (8), Exposure-$Foreign_{f,pre}$ and Exposure-$Local_{f,pre}$ are the average of FX-Foreign Inflows $_{f,yq}$ and of FX-Local Inflows $_{f,yq}$ in the period from 2005:Q1 to 2007:Q1, respectively. $Post_{yq}$ is a dummy with value 1 from 2007:Q2 to 2008:Q2 and 0 from 2006:Q1 to 2007:Q1. Firm Controls include $ROA_{f,y-1}$, $Size_{f,y-1}$, $Imports_{f,yq-1}$, $Exports_{f,yq-1}$. $Default_{f,yq}$ and $Relationships_{f,yq}$. Bank Controls include: $BankCET1_{f,yq-1}$; $BankROA_{f,yq-1}$; $BankSIZE_{f,yq-1}$; $BankNPL_{f,yq-1}$; $BankSaving_{f,yq-1}$; $BankChecking_{f,yq-1}$; BankFX-$Funds_{f,yq-1}$. Both Bank and Firm controls are fully interacted with the $Post_{yq}$ dummy. The sign "-" denotes cases where a variable (or a group of variables or fixed effects) is spanned out by other controls/fixed effects. Standard errors in parentheses are clustered at the firm level. *** p<0.01, ** p<0.05, *p<0.1.

TABLE 8: Real effects – Capital Controls and Trade during the Boom and the Bust

Panel A: Baseline results for exports and imports

	(1)	(2)	(3)	(4)
	$Exports_{f,yq}$	$Imports_{f,yq}$	$Exports_{f,yq}$	$Imports_{f,yq}$
$Crisis_{yq}$*Exposure-$Foreign_{f,pre}$	5.2213^{***}	-1.7723	2.3819^{*}	-0.6226
	(1.814)	(1.722)	(1.231)	(1.094)
$Crisis_{yq}$*Exposure-$Local_{f,pre}$	-1.1536	2.7480	-2.1527	1.8147
	(3.439)	(2.032)	(2.397)	(1.552)
$Post_{yq}$ *Exposure-$Foreign_{f,pre}$	-1.0216	-3.1762^{**}	1.5957	-3.1905^{***}
	(2.254)	(1.255)	(1.508)	(0.994)
$Post_{yq}$ *Exposure-$Local_{f,pre}$	-1.4590	0.9796	-1.2024	0.5269
	(3.349)	(1.634)	(2.327)	(1.311)
N	15269	25294	25391	37484
R^2	0.8476	0.8396	0.8747	0.8534
Firm Controls*[Post ; Crisis]	YES	YES	YES	YES
Bank Controls*[Post ; Crisis]	YES	YES	YES	YES
Firm FE	YES	YES	YES	YES
Industry*Year-quarter FE	YES	YES	YES	YES
Companies Active in Both	Excluded	Excluded	Included	Included

Panel B: Mechanism – Growth of total liabilities

	(1) $Exports_{f,yq}$	(2) $Imports_{f,yq}$	(3) $Exports_{f,yq}$	(4) $Imports_{f,yq}$
$Crisis_{yq}*(-\Delta_{1y}Liabilities_{f,2007}^{predicted})$	4.8597***	-1.9240	2.0079*	-0.7038
	(1.716)	(1.582)	(1.127)	(0.985)
$Crisis_{yq}*Exposure\text{-}Local_{f,pre}$	-1.0676	2.8161	-2.2189	1.8829
	(3.454)	(2.025)	(2.399)	(1.551)
$Crisis_{yq}*(-\Delta_{1y}Liabilities_{f,2007}^{residual})$	0.1591	0.0660	0.0524	0.0748
	(0.123)	(0.094)	(0.084)	(0.075)
$Post_{yq}*(-\Delta_{1y}Liabilities_{f,2007}^{predicted})$	-1.7336	-2.9384**	1.0659	-2.7399***
	(1.931)	(1.171)	(1.297)	(0.881)
$Post_{yq}*Exposure\text{-}Local_{f,pre}$	-1.6858	0.7582	-1.0353	0.5738
	(3.359)	(1.594)	(2.311)	(1.275)
$Post_{yq}*(-\Delta_{1y}Liabilities_{f,2007}^{residual})$	-0.3648**	-0.3850***	-0.3002**	-0.4039***
	(0.153)	(0.093)	(0.127)	(0.076)
N	14998	24868	25091	37016
R^2	0.8481	0.8401	0.8751	0.8538
Firm Controls*[Post ; Crisis]	YES	YES	YES	YES
Bank Controls*[Post ; Crisis]	YES	YES	YES	YES
Firm FE	YES	YES	YES	YES
Industry*Year-Quarter FE	YES	YES	YES	YES
Companies active in Both	Excluded	Excluded	Included	Included

Panel C: Exports – Companies sorted according to proxies of financing constraints

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
					$Exports_{f,yq}$				
	Loan Interest Rate			% Collateralized Debt			% Short-Term Debt (≤1y)		
	All	Low	High	All	Low	High	All	Low	High
$Crisis_{yq}$*$Exposure\text{-}Foreign_{f,pre}$	5.1960***	3.2374	7.5520***	5.4803***	4.4204**	20.6088***	4.9244***	0.2981	9.5632***
	(1.882)	(2.384)	(2.495)	(1.903)	(2.127)	(7.484)	(1.852)	(3.074)	(3.457)
$Crisis_{yq}$*$Exposure\text{-}Local_{f,pre}$	-1.9590	1.1328	-6.1956	-1.3611	-3.6624	8.0004*	-1.0751	1.2764	-0.9220
	(3.659)	(4.816)	(4.691)	(3.550)	(4.843)	(4.521)	(3.551)	(6.815)	(4.150)
$Post_{yq}$ *$Exposure\text{-}Foreign_{f,pre}$	-0.8325	1.8080	-5.1024	-0.8090	-0.7187	-1.6442	-1.0068	-0.2588	0.5760
	(2.319)	(2.468)	(3.401)	(2.319)	(2.350)	(8.452)	(2.300)	(3.701)	(3.375)
$Post_{yq}$ *$Exposure\text{-}Local_{f,pre}$	0.3068	-0.5093	1.3247	-1.5242	-2.0529	-3.6530	-1.6852	1.7283	-3.9347
	(3.247)	(4.388)	(3.135)	(3.495)	(4.906)	(3.878)	(3.451)	(7.237)	(3.959)
N	14172	7103	7069	14162	7151	7011	14269	7176	7093
R^2	0.8489	0.8743	0.8386	0.8440	0.8637	0.8453	0.8477	0.8552	0.8635
Firm Controls*[Post ; Crisis]	YES	YES	YES	YES	YES	YES	YES	YES	YES
Bank Controls*[Post ; Crisis]	YES	YES	YES	YES	YES	YES	YES	YES	YES
Firm FE	YES	YES	YES	YES	YES	YES	YES	YES	YES
Industry*Time FE	YES	YES	YES	YES	YES	YES	YES	YES	YES
Companies active in Both	Excluded	Excluded	Excluded	Excluded	Excluded	Excluded	Excluded	Excluded	Excluded

Panel D: Imports – Companies sorted according to proxies of financing constraints

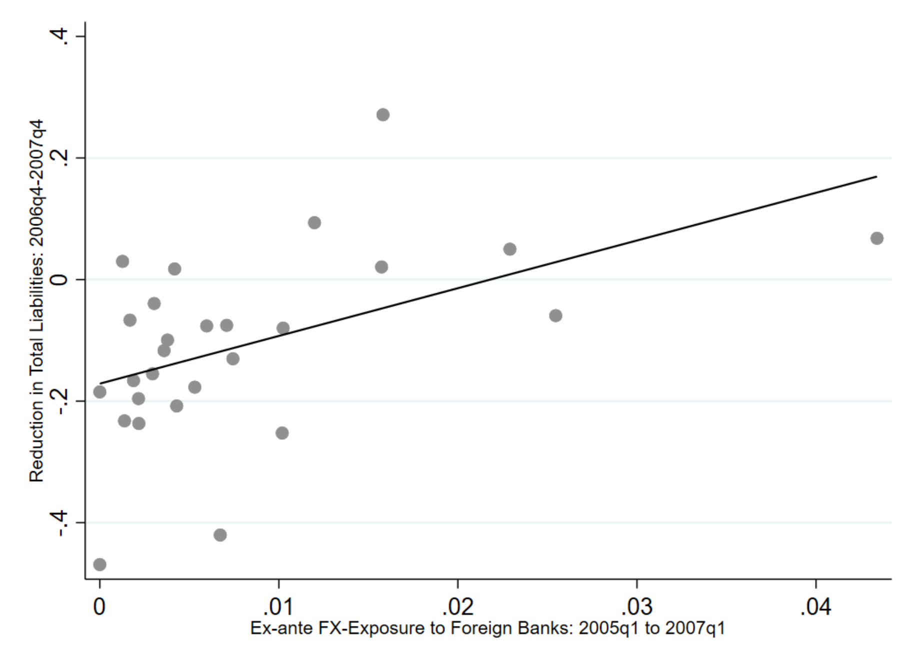
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	$Imports_{f,yq}$								
	Loan Interest Rate			% Collateralized Debt			% Short-Term Debt (≤1y)		
	All	Low	High	All	Low	High	All	Low	High
$Crisis_{yq}$*Exposure-$Foreign_{f,pre}$	-0.7940	-0.0600	-2.2971	-1.0514	0.1316	-2.1827	-1.4203	-1.5812	-0.7398
	(1.614)	(2.119)	(2.281)	(1.632)	(1.915)	(4.538)	(1.709)	(2.218)	(2.728)
$Crisis_{yq}$*Exposure-$Local_{f,pre}$	2.8681	4.0890	-0.5642	2.8238	-0.7546	11.3877**	2.9607	7.1357*	4.1200*
	(2.077)	(2.897)	(2.911)	(2.056)	(1.925)	(4.783)	(2.053)	(4.157)	(2.434)
$Post_{yq}$ *Exposure-$Foreign_{f,pre}$	-3.1068**	-3.9203*	-2.6509*	-3.5807***	-2.3797	-8.5667**	-3.3384***	-3.9725**	-3.4836**
	(1.256)	(2.170)	(1.488)	(1.250)	(1.551)	(3.841)	(1.256)	(1.985)	(1.770)
$Post_{yq}$ *Exposure-$Local_{f,pre}$	1.0180	1.3016	-1.5710	1.2012	1.7479	-1.3561	1.0019	-2.4461	2.1644
	(1.676)	(2.024)	(2.901)	(1.660)	(1.906)	(3.740)	(1.655)	(3.848)	(1.984)
N	24063	12017	12046	24171	12138	12033	24022	12119	11903
R^2	0.8389	0.8426	0.8514	0.8376	0.8620	0.8262	0.8366	0.8461	0.8425
Firm Controls*[Post ; Crisis]	YES	YES	YES	YES	YES	YES	YES	YES	YES
Bank Controls*[Post ; Crisis]	YES	YES	YES	YES	YES	YES	YES	YES	YES
Firm FE	YES	YES	YES	YES	YES	YES	YES	YES	YES
Industry*Time FE	YES	YES	YES	YES	YES	YES	YES	YES	YES
Companies active in Both	Excluded	Excluded	Excluded	Excluded	Excluded	Excluded	Excluded	Excluded	Excluded

This table shows the impact of capital controls on firm-level trade. In Panel A, we report how exposure to local and foreign banks affect exports and imports, during capital controls (boom) and during the GFC (bust). The dependent variable is either the logarithm of (1 + exports), $Exports_{f,yq}$, or of (1+imports), $Imports_{f,yq}$ of firm f in year-quarter yq. Exposure-$Foreign_{f,pre}$ and Exposure-$Local_{f,pre}$ are the average of FX-Foreign $Inflows_{f,yq}$ and of FX-Local $Inflows_{f,yq}$ in the period from 2005:Q1 to 2007:Q1, respectively. $Post_{yq}$ is a dummy with value 0 (1) from 2006:Q1 to 2007:Q1 (2007:Q2 to 2009:Q4). $Crisis_{yq}$ is a dummy with value 1 from 2008:Q3 to 2009:Q4 and 0 before. Firm Controls include $ROA_{f,y-1}$, $Size_{f,y-1}$ and $Imports_{f,yq-1}$ ($Exports_{f,yq-1}$) in regressions where exports (imports) is the dependent variable. Bank Controls include: $BankCET1_{f,yq-1}$; $BankROA_{f,yq-1}$; $BankSIZE_{f,yq-1}$; $BankNPL_{f,yq-1}$; $BankSaving_{f,yq-1}$; $BankChecking_{f,yq-1}$; BankFX-$Funds_{f,yq-1}$. Both Bank and Firm controls are fully interacted with the $Post_{yq}$ and $Crisis_{yq}$ dummies. In Panel B, we replicate panel A, replacing Exposure-$Foreign_{f,pre}$ with -$\Delta_{1y}Liabilities_{f,2007}^{predicted}$, the yearly reduction in total liabilities that it predicts in 2007 (in a cross-sectional regression with industry fixed effects – coefficient is equal to 1.1397, significance at 1% level). We also include the residual heterogeneity, denoted by -$\Delta_{1y}Liabilities_{f,2007}^{residual}$. In panels C and D, respectively, we repeat the same exercise for exports and imports, sorting companies based on proxies of financial constraints - i.e. indicators of high interest rate, collateral requirements and percentage of short-term debt (maturity smaller or equal than 1 year). These are taken as weighted average of related variables from the credit registry (after taking out bank*industry*year-quarter fixed effects) - with weights given by the loan share over total bank debt – over the period 2005:Q1-2007:Q1. A company is defined as High (Low) Interest Rate/% Collateralized Debt/% Short-Term Debt if its value is above (below) the median in the regression sample.

Appendix

FIGURE A1: Ex-ante FX-Exposures and Reduction in Total Liabilities : Industry-level

Panel A: FX-Exposure to Foreign Banks



Panel B: FX-Exposure to Local Banks

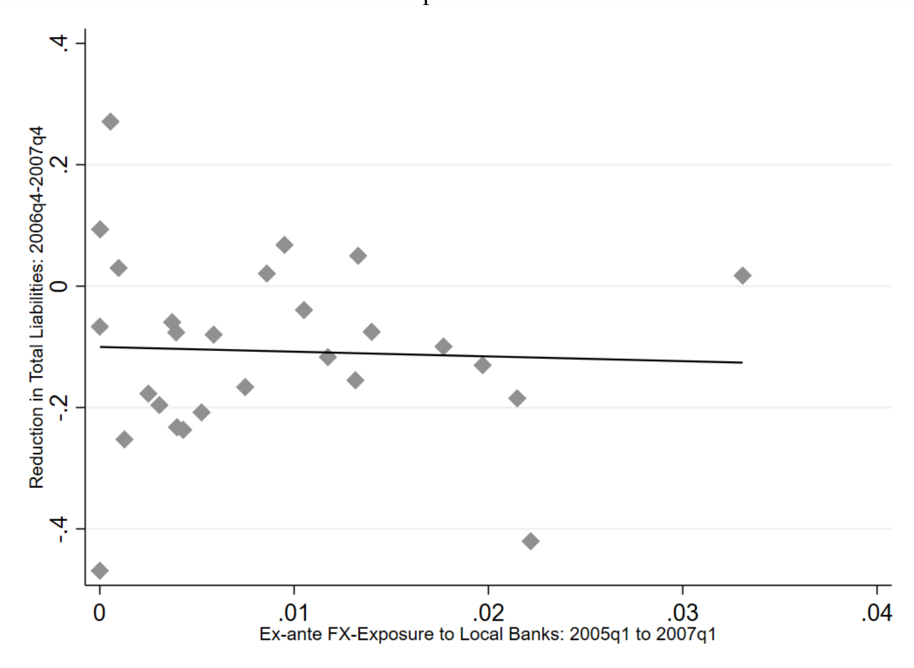


TABLE A1: Other Summary Statistics - Macro and Industrial Level Variables (2006:Q1-2007:Q1)

VARIABLES	N	Mean	P25	P50	P75	SD
Firm-level variables (2006:Q1-2007:Q1)						
$ExposureLiab_{f,pre}$	14,125	0.0244	0.000517	0.00702	0.0319	0.0408
$Exposure_{f,2007:Q1}$	14,125	0.0124	0	0	0.00534	0.0341
$AvgLogExposure_{f,pre}$	14,125	1.827	0.310	0.970	2.804	2.008
$Exposure_{f,2005}$	14,125	0.0127	0	0.00113	0.01419	0.0208
$Exposure\text{-}Foreign\text{-}Liab_{f,pre}$	5,751	0.0253	0.00237	0.00937	0.0302	0.0423
$Exposure\text{-}Foreign_{f,2007:Q1}$	5,751	0.0124	0	0	0.00364	0.0379
$Exposure\text{-}Foreign\text{-}Log_{f,pre}$	5,751	1.894	0.568	1.171	2.581	1.825
$Exposure\text{-}Foreign_{f,2005}$	5,751	0.0162	0	0.00402	0.0160	0.0352
$Exposure\text{-}Local\text{-}Liab_{f,pre}$	12,176	0.0164	0.000240	0.00315	0.0190	0.0288
$Exposure\text{-}Local_{f,2007:Q1}$	12,176	0.00836	0	0	0.00120	0.0234
$Exposure\text{-}Local\text{-}Log_{f,pre}$	12,176	1.428	0.204	0.675	2.145	1.655
$Exposure\text{-}Local_{f,2005}$	12,176	0.00714	0	0.00017	0.00782	0.0142
$-\Delta Liabilities_{f,2007}^{predicted}$	5,433	0.00735	0.0157	0.00554	0.01638	.02843
$-\Delta Liabilities_{f,2007}^{residual}$	11,597	0.00475	-0.1662	0.01735	0.18658	0.34617
Macroeconomic Variables (2006:Q1-2008:Q2)						
Δi_{yq-1}	10	0.0105	-0.00267	0.0168	0.0198	0.0133
$\Delta \pi_{yq-1}$	10	0.0630	0.0572	0.0619	0.0763	0.0138
ΔGDP_{yq-1}	10	0.0504	0.0448	0.0494	0.0577	0.00766
ΔVIX_{yq-1}	10	0.184	-0.0432	0.149	0.407	0.289
Δe_{yq-1}	10	-0.0624	-0.1175	-0.0469	-0.0077	0.0903
(Continued below)						
Industry-Level Variables (2006:Q1-2007:Q1)						
$Employment_{i,yq}$	135	4.547	4.486	4.611	4.732	0.380
$Exposure\text{-}Foreign_{i,pre}$	135	0.00817	0.00216	0.00430	0.0102	0.00950
$Exposure\text{-}Local_{i,pre}$	135	0.00881	0.00250	0.00586	0.0133	0.00822
$Exposure\text{-}Foreign\text{-}Liab_{i,pre}$	135	0.0165	0.00339	0.00968	0.0259	0.0166
$Exposure\text{-}Local\text{-}Liab_{i,pre}$	135	0.0208	0.00548	0.0145	0.0269	0.0210
$Exposure\text{-}Foreign\text{-}Log_{i,pre}$	135	1.607	0.594	1.437	2.433	1.244
$Exposure\text{-}Foreign\text{-}Log_{i,pre}$	135	1.956	0.786	1.721	3.045	1.321

$Size_{i,yq-1}$	135	8.764	8.059	8.540	9.036	1.017
$ROA_{i,yq-1}$	135	0.0329	0.0187	0.0345	0.0555	0.0275
$Imports_{i,yq-1}$	135	6.543	5.569	6.752	7.892	1.799
$Exports_{i,yq-1}$	135	5.630	4.174	6.274	7.243	2.221

Firm-level Variables. $ExposureLiab_{f,pre}$ is the average of the ratio between FX-debt flows and total liabilities over the period 2005:Q1-2007:Q1. $Exposure_{f,2007:Q1}$ is the ratio between FX-debt flows and total assets as of 2007:Q1. $AvgLogExposure_{f,pre}$ is the average of the logarithm of (1 + FX-debt flow) during the period 2005:Q1-2007:Q1. $Exposure_{f,2005}$ is the average of the ratio between FX-debt flows and total assets over the period 2005:Q1-2005:Q4. Exposure-Foreign-$Liab_{i,pre}$ is the average of the ratio between FX-debt flows from foreign banks and total liabilities over the period 2005:Q1-2007:Q1. Exposure-$Foreign_{f,2007:Q1}$ is the ratio between FX-debt flows from foreign banks and total assets as of 2007:Q1. Exposure-Foreign-$Log_{f,pre}$ is the average of the logarithm of (1 + FX-debt flow from foreign banks during the period 2005:Q1-2007:Q1). Exposure-Local-$Liab_{i,pre}$ is the average of the ratio between FX-debt flows from local banks and total liabilities over the period 2005:Q1-2007:Q1. Exposure-$Local_{f,2007:Q1}$ is the ratio between FX-debt flows from local banks and total assets as of 2007:Q1. Exposure-Local-$Log_{f,pre}$ is the average of the logarithm of (1 + FC-debt flow from local banks during the period 2005:Q1-2007:Q1). $-\Delta_{1y}Liabilities_{f,2007}^{predicted}$ is the yearly reduction in total liabilities predicted by Exposure-$Foreign_{f,pre}$ in a cross-sectional regression in 2007 with industry fixed effects. Its summary statistics are computed over companies ex-ante active in foreign FX-debt markets (for all others, the value is constant and equal to 0). The residual heterogeneity in total liabilities from same regression is $-\Delta_{1y}Liabilities_{f,2007}^{residual}$. Macroeconomic Variables (2006:Q1-2008:Q2). Δi_{yq-1} is the lagged yearly growth of the interbank rate. $\Delta\pi_{yq-1}$ is the lagged yearly inflation rate. ΔGDP_{yq-1} is the lagged yearly growth rate of GDP. ΔVIX_{yq-1} is the lagged yearly growth rate of VIX. Δe_{yq-1} is the lagged yearly growth rate of the exchange rate – defined as Colombian pesos per 1US$. Industry-Level Variables (2006:Q1-2007:Q1). $Employment_{i,yq}$ is the logarithm of the employment index. The following exposure measures are retrieved as weighted averages of firm-level correspondent variables. Weights are given by the ratio of a company's total assets to total industrial assets, as of the end of 2006. Exposure-$Foreign_{i,pre}$ is the industry-level weighted average of firm-level FX-exposure to foreign banks, rescaled by total assets. Exposure-$Local_{i,pre}$ is the industry-level weighted average of firm-level FX-exposure to local banks, rescaled by total assets. Exposure-Foreign-$Liab_{i,pre}$ is the industry-level weighted average of firm-level FX-exposure to foreign banks, rescaled by total liabilities. Exposure-Local-$Liab_{i,pre}$ is the industry-level weighted average of firm-level FX-exposure to local banks, rescaled by total liabilities. Exposure-Foreign-$Log_{i,pre}$ is the industry-level weighted average of firm-level FX-exposure to foreign banks, defined in logs. Exposure-Local-$Log_{i,pre}$ is the industry-level weighted average of firm-level FX-exposure to local banks, defined in logs. The remaining variables are defined as weighted averages of firm-level correspondent variables. Weights are given by the time-varying ratio of a company's total assets to total industrial assets. $Size_{i,yq-1}$ is the lagged average of firm log(assets). $ROA_{i,yq-1}$ is the lagged average firm ROA. $Imports_{i,yq-1}$ is the lagged average of log-firm imports. $Exports_{i,yq-1}$ is the lagged average of log-firm exports.

TABLE A2: Impact of Capital Controls on FX-Debt Inflows – Robustness Checks

Panel A: Unconditional impact across market segments

	(1)	(2)	(3)
$Post_{yq}$	-0.002**	-0.006*	-0.010***
	(0.001)	(0.003)	(0.003)
N	16741	7622	3925
R^2	0.4044	0.3746	0.3696
Companies	Local	Both	Foreign
Quarter FE	YES	YES	YES
Macro Controls	YES	YES	YES
Firm Controls	YES	YES	YES
Bank Controls	YES	YES	YES

Panel B: Conditional impact on different time windows around the policy shock

	(1)	(2)	(3)	(4)	(5)
	2007:Q1-2007:Q2	2006:Q4-2007:Q3	2006:Q3-2007:Q4	2006:Q2-2008:Q1	2006:Q1-2008:Q2
$Post_{yq}$ *$Exposure_{f,pre}$	-0.3054***	-0.3930***	-0.3984***	-0.5038***	-0.4609***
	(0.089)	(0.070)	(0.065)	(0.059)	(0.051)
N	5636	11327	16980	22650	28288
R^2	0.7071	0.5357	0.4999	0.4850	0.4615
Firm FE	YES	YES	YES	YES	YES
Industry*Year-quarter FE	YES	YES	YES	YES	YES
Firm Controls*Post	YES	YES	YES	YES	YES
Bank Controls*Post	YES	YES	YES	YES	YES

Panel C: Conditional Impact - Alternative Definitions of Exposure

	(1)	(2)	(3)	(4)
Post*ExposureLiab$_{f,pre}$	-0.2447***			
	(0.027)			
Post*Exposure$_{f,2007:Q1}$		-0.2119***		
		(0.036)		
Post*AvgLogExposure$_{f,pre}$			-0.0040***	
			(0.000)	
Post*Exposure$_{f,2005}$				-.1675***
				(0.031)
N	28288	28288	28288	28288
R^2	0.4590	0.4525	0.4497	0.4464
Firm FE	YES	YES	YES	YES
Industry*Year-quarter FE	YES	YES	YES	YES
Firm Controls*Post	YES	YES	YES	YES
Bank Controls*Post	YES	YES	YES	YES

The dependent variable is FX Inflows$_{f,yq}$. Panel A shows the effect of the introduction of capital controls on total FX debt inflows for firms borrowing in FX from local intermediaries (column 1), both local intermediaries and foreign (column 2), and foreign only (column 3). Panel B shows the effect of the introduction of capital controls on total FX debt inflows in different symmetric time-windows around the introduction of capital controls in 2007:Q2. Panel C shows the effect of the introduction of capital controls on total FX debt inflows, depending on different definitions of pre-policy exposure to FX debt inflows. Post$_{yq}$ is a dummy with value 1 (0) from 2007:Q2 to 2008:Q2 (2006:Q1 to 2007:Q1). ExposureLiab$_{f,pre}$ is the average of the ratio between FX debt inflows and total liabilities from 2005:Q1 to 2007:Q1. Exposure$_{f,2007:Q1}$ is the dependent variable as of 2007:Q1. AvgLogExposure$_{f,pre}$ is the average log FX debt inflows in the period from 2005:Q1 to 2007:Q1. Exposure$_{f,2005}$ is the average FX-debt inflow rescaled by total assets between 2005:Q1 and 2005:Q4. Macro Controls include lagged: GDP growth rate; inflation rate and log(VIX). Firm Controls include ROA$_{f,y-1}$, Size$_{f,y-1}$, Imports$_{f,yq-1}$, Exports$_{f,yq-1}$. Bank Controls include: BankCET1$_{f,yq-1}$; BankROA$_{f,yq-1}$; BankSIZE$_{f,yq-1}$; BankNPL$_{f,yq-1}$; BankSaving$_{f,yq-1}$; BankChecking$_{f,yq-1}$; BankFX$_{f,yq-1}$. Standard errors in parentheses are double-clustered at the firm and industry*year-quarter level.*** p<0.01, ** p<0.05, *p<0.1.

TABLE A3: Impact of Capital Controls on FX-Debt Inflows – Oster Test

	(1) $\tilde{R}^2 = 1.3\,\hat{R}^2$	(2) $\tilde{R}^2 = 1$
$\tilde{\delta}$	4.712	1.350

This table shows the robustness of our estimates in Table 3 to the Oster (2017) test for selection into the treatment along unobservables. In column (1), the coefficient of proportionality $\tilde{\delta}$ is estimated under the assumptions that the maximum R-squared is equal to 1.3 $\hat{R}^2$, where $\hat{R}^2$ is the R-squared reported in column (7) of Table 3. In column (2), the maximum R-squared is assumed to be equal to 1. Note: the baseline version of the model only includes the full interaction of the $Post_{yq}$ dummy with $Exposure_{f,pre}$. The test refers to the stability of the coefficient for $Post_{yq}$*$Exposure_{f,pre}$.

TABLE A4: Substitution of FX with Peso Debt: Intensive Margin - Oster Test

	(1) $\tilde{R}^2=\tilde{R}^2_{ft}$	(2) $\tilde{R}^2=1$
Quantity		
$\tilde{\delta}^*_{Post*Both}$	8.843	2.618
$\tilde{\delta}^*_{Post*Foreign}$	17.23	5.117
Price		
$\tilde{\delta}^*_{Post*Both}$	-1.343	-0.386
$\tilde{\delta}^*_{Post*Foreign}$	-7.866	-2.263

This table shows the robustness of our estimates in Tables 5 to the Oster (2017) test for selection into the treatment along unobservables. In column (1), the coefficient of proportionality $\tilde{\delta}$ is estimated under the assumptions that the maximum R-squared is equal to the R-square obtained by saturating the model with firm*bank, firm*year-quarter and bank*year-quarter fixed effects. In column (2), the maximum R-squared is assumed to be equal to 1. Note: the baseline version of the model only includes the full interaction of the $Post_{yq}$ dummy with the $Foreign_{f,pre}$ and $Both_{f,pre}$ dummies, respectively. The tests refer to the stability of the coefficient for $Post_{yq}$*$Both_{f,pre}$ and $Post_{yq}$*$Foreign_{f,pre}$, respectively, compared in the baseline version of the model and in one including firm*bank, bank*year-quarter fixed effects and firm controls interacted with the $Post_{yq}$ dummy.

TABLE A5: Substitution with Peso Debt from Local Banks - Collapsed Pre-Post Time Dimension

	(1) Peso Loan$_{f,b,yq}$	(2) Interest Rate$_{f,b,yq}$
Post$_{yq}$*Both$_{f,pre}$	-0.002	0.175
	(0.037)	(0.142)
Post$_{yq}$*Foreign$_{f,pre}$	-0.103**	0.478**
	(0.045)	(0.195)
N	17074	17074
R^2	0.913	0.823
Firm Controls*Post	YES	YES
Firm*Bank FE	YES	YES
Bank*Year-quarter FE	YES	YES
Industry*Year-quarter FE	YES	YES
Loan Controls*Post	YES	YES

This table shows the effect of capital controls on the quantity and price of commercial (peso) credit granted from Colombian banks. The baseline category is given by companies borrowing in FX before 2007:Q2 from local banks only. In column (1), the dependent variable is formally defined as the logarithm of the mean of (1+stock of peso debt provided by bank b to firm f) in the pre-period (2006:Q1 to 2007:Q1) and the post-period (2007:Q2 to 2008:Q2). In column (2), the dependent variable is formally defined as the mean of the interest rate applied on debt provided by bank b to firm f in the pre-period (2006:Q1 to 2007:Q1) and the post period (2007:Q2 to 2008:Q2). Equally, independent variables are mean-collapsed in the pre-period (2006:Q1 to 2007:Q1) and the post period (2007:Q2 to 2008:Q2). Foreign$_{f,pre}$ is a dummy with value 1 if a company borrowed in FX only from foreign banks before 2007:Q2 and 0 otherwise. Both$_{f,pre}$ refers to companies resorting to both local and foreign banks for FX credit before 2007:Q2. Post$_{yq}$ is a dummy with value 1 from 2007:Q2 to 2008:Q2 and 0 from 2006:Q1 to 2007:Q1. Firm Controls include ROA$_{f,y-1}$, Size$_{f,y-1}$, Imports$_{f,yq-1}$, Exports$_{f,yq-1.}$ Default$_{f,yq}$ and Relationships$_{f,yq.}$ Loan Controls include: Maturity$_{f,b,yq}$ and Collateral$_{f,b,yq}$. Both Firm and Loan controls are fully interacted with the Post$_{yq}$ dummy. Standard errors in parentheses are clustered at the firm level.*** p<0.01, ** p<0.05, *p<0.1.

TABLE A6: Substitution with Peso Debt from Local Banks: Impact Conditional on Pre-policy FX Exposure

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	Peso Loan$_{f,b,yq}$					Interest Ratef,$_{b,yq}$				
Post$_{yq}$*Exposure-Foreign$_{f,pre}$	-3.035**	-3.159*	-2.014	-1.846	-2.007*	2.990	8.888***	6.208*	9.936***	9.103**
	(1.235)	(1.761)	(1.448)	(1.421)	(1.134)	(3.198)	(3.433)	(3.501)	(3.357)	(3.552)
Post$_{yq}$*Exposure-Local$_{f,pre}$	4.383***	3.700***	4.382***	4.291***	3.793***	-36.107***	-27.055***	-33.134***	-30.305***	-30.710***
	(1.313)	(1.422)	(1.248)	(1.247)	(1.199)	(3.999)	(3.984)	(4.116)	(3.973)	(4.135)
N	102035	102035	102035	102035	102035	102035	102035	102035	102035	102035
R^2	0.005	0.262	0.789	0.791	0.802	0.067	0.109	0.537	0.609	0.625
Firm Controls*Post	NO	YES	YES	YES	YES	NO	YES	YES	YES	YES
Firm*Bank FE	NO	NO	YES	YES	YES	NO	NO	YES	YES	YES
Bank*Year-Quarter FE	NO	NO	NO	YES	YES	NO	NO	NO	YES	YES
Industry*Year-Quarter FE	NO	NO	NO	NO	YES	NO	NO	NO	NO	YES
Loan Controls*Post	NO	NO	NO	NO	YES	NO	NO	NO	NO	YES

This table shows the effect of capital controls on the quantity (columns 1-5) and price (columns 6-10) of commercial (peso) credit granted from Colombian banks. The dependent variable is defined as the logarithm of the loan in pesos granted from bank b to firm f in year-quarter yq or as the interest rate (in percentage points) applied over the same loans. Exposure-Foreign$_{f,pre}$ and Exposure-Local$_{f,pre}$ are the average of FX-Foreign Inflows $_{f,yq}$ and of FX-Local Inflows $_{f,yq}$ in the period from 2005:Q1 to 2007:Q1, respectively. Firm Controls include ROA$_{f,y-1}$, Size$_{f,y-1}$, Imports$_{f,yq-1}$, Exports$_{f,yq-1,}$ Default$_{f,yq}$ and Relationships$_{f,yq.}$ Loan Controls include: Maturity$_{f,b,yq}$ and Collateral$_{f,b,yq}$. Both Firm and Loan controls are eventually fully interacted with the Post$_{yq}$ dummy. The sign "-" denotes cases where a variable (or a group of variables or of fixed effects) is spanned out by other controls and/or fixed effects. Standard errors in parentheses are clustered at the firm level. *** p<0.01, ** p<0.05, *p<0.1.

TABLE A7: Substitution of FX with Peso Debt - Intensive Margin: Impact Conditional on Pre-policy FX Exposure – Different Definitions

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	$PesoLoan_{f,b,yq}$				$InterestRate_{f,b,yq}$			
$Post_{yq}$*Exposure-Foreign-$Liab_{f,pre}$	-1.1788*				4.6610**			
	(0.668)				(2.181)			
$Post_{yq}$*Exposure-Local-$Liab_{,pre}$	1.8453***				-17.0125***			
	(0.715)				(2.399)			
$Post_{yq}$*Exposure-$Foreign_{f,2007:Q1}$		-1.4477**				6.3302***		
		(0.673)				(2.375)		
$Post_{yq}$*Exposure- $Local_{f,2007:Q1}$		1.3919*				-12.1667***		
		(0.768)				(2.689)		
$Post_{yq}$*Exposure-Foreign-$Log_{f,pre}$			-0.0204*				0.1279***	
			(0.011)				(0.038)	
$Post_{yq}$*Exposure-Local-$Log_{f,pre}$			0.0296***				-0.2563***	
			(0.011)				(0.037)	
$Post_{yq}$*Exposure- $Foreign_{f,2005}$				-1.6130**				3.7388*
				(0.810)				(2.248)
$Post_{yq}$*Exposure- $Local_{f,2005}$				2.6620**				-25.8490***
				(1.190)				(4.033)
N	102035	102035	102035	102035	102035	102035	102035	102035
R^2	0.8019	0.8019	0.8019	0.8019	0.6249	0.6245	0.6248	0.6248
Firm Controls*Post	YES	YES	YES	YES	YES	YES	YES	YES
Firm*Bank FE	YES	YES	YES	YES	YES	YES	YES	YES
Bank*Year-Quarter FE	YES	YES	YES	YES	YES	YES	YES	YES
Industry*Year-Quarter FE	YES	YES	YES	YES	YES	YES	YES	YES
Loan Controls*Post	YES	YES	YES	YES	YES	YES	YES	YES

This table shows the effect of capital controls on the quantity (columns 1-4) and price (columns 1-8) of peso-credit granted to companies from Colombian banks, depending on a firm's pre-policy FX-exposure to foreign and local banks, respectively. $Post_{yq}$ is a dummy with value 1(0) from 2007:Q2 to 2008:Q2 (2006:Q1 to 2007:Q1). In columns (1) and (5), Exposure-Foreign-$Liab_{f,pre}$ and Exposure-Local-$Liab_{f,pre}$ are the average firm-level FX debt inflows from foreign and local banks in the period 2005:Q1-2007:Q1, rescaled by total liabilities. In columns (2) and (6), Exposure-$Foreign_{f,2007:Q1}$ and Exposure-$Local_{f,2007:Q1}$ are given by the 2007:Q1 firm-level values of foreign and local FX-debt inflows over total assets. In columns (3) and (7), Exposure-Foreign-$Log_{f,pre}$ and Exposure-Local-$Log_{f,pre}$ are the average firm-level log FX debt inflows from foreign and local banks in the period 2005:Q1- 2007:Q1. In columns (4) and (8), Exposure-$Foreign_{f,2005}$ and Exposure-$Local_{f,2005}$ represent the average firm-level FX-debt inflow (rescaled by total assets) from local and foreign banks over the period 2005:Q1 to 2005:Q4. Firm Controls include $ROA_{f,y-1}$, $Size_{f,y-1}$, $Imports_{f,yq-1}$, $Exports_{f,yq-1.}$ $Default_{f,yq}$ and $Relationships_{f,yq.}$ Loan Controls include: $Maturity_{f,b,yq}$ and $Collateralized_{f,b,yq}$. Each regression includes Firm and Loan controls, fully interacted with the $Post_{yq}$ dummy and firm*bank, bank*year-quarter and industry*year-quarter fixed effects. Standard errors in parentheses are clustered at the firm level.*** p<0.01, ** p<0.05, *p<0.1.

TABLE A8: Real Effects – Capital Controls and Trade during the Boom and the Bust – Robustness Checks

Panel A: Different Specifications of the Model - Exports

	(1)	(2)	(3)	(4)
		$Exports_{f,yq}$		
$Crisis_{yq}$ *$Exposure\text{-}Foreign_{f,pre}$	4.9321**	4.5420**	5.1335***	5.2213***
	(2.213)	(2.092)	(1.617)	(1.814)
$Crisis_{yq}$*$Exposure\text{-}Local_{f,pre}$	2.4782	-0.7784	-1.0380	-1.1536
	(4.900)	(4.699)	(2.893)	(3.439)
$Post_{yq}$ *$Exposure\text{-}Foreign_{f,pre}$	-0.6415	-1.0928	-0.2990	-1.0216
	(2.717)	(2.486)	(2.388)	(2.254)
$Post_{yq}$*$Exposure\text{-}Local_{f,pre}$	-0.7232	-3.6735	-2.1139	-1.4590
	(5.006)	(4.401)	(2.646)	(3.349)
N	15269	15269	15269	15269
R^2	0.0019	0.1015	0.8173	0.8476
Firm Controls*[Post; Crisis]	NO	YES	YES	YES
Bank Controls*[Post; Crisis]	NO	YES	YES	YES
Firm FE	NO	NO	YES	YES
Industry*Year-quarter FE	NO	NO	NO	YES
Companies active in both	Excluded	Excluded	Excluded	Excluded

Panel B: Different specifications of the model - Imports

	(1)	(2)	(3)	(4)
		Imports$_{f,yq}$		
Crisis$_{yq}$ *Exposure-Foreign$_{f,pre}$	-1.7073	-0.6695	-1.2383	-1.7723
	(1.938)	(2.670)	(2.030)	(1.722)
Crisis$_{yq}$*Exposure-Local$_{f,pre}$	2.9436	1.9500	2.2959	2.7480
	(2.724)	(2.792)	(1.781)	(2.032)
Post$_{yq}$ *Exposure-Foreign$_{f,pre}$	-5.1875***	-4.9645***	-4.7530***	-3.1762**
	(1.323)	(1.758)	(1.351)	(1.255)
Policy$_{yq}$*Exposure-Local$_{f,pre}$	0.6495	0.1407	0.3781	0.9796
	(2.502)	(2.528)	(1.305)	(1.634)
N	25294	25294	25294	25294
R^2	0.0705	0.2629	0.8166	0.8396
Firm Controls*[Post; Crisis]	NO	YES	YES	YES
Bank Controls*[Post; Crisis]	NO	YES	YES	YES
Firm FE	NO	NO	YES	YES
Industry*Year-quarter FE	NO	NO	NO	YES
Companies active in both	Excluded	Excluded	Excluded	Excluded

Panel C: Oster Test – Imports and Exports

	$\widetilde{R^2} = 1$
Imports	
$\tilde{\delta}_{Post*Exposure\text{-}Foreign}$	5.38
Exports	
$\tilde{\delta}_{Crisis*Exposure\text{-}Foreign}$	32.97

Panel D: Different Definitions of the Exposure variables – Imports and Exports

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Exports$_{f,yq}$				Imports$_{f,yq}$			
Crisis$_{yq}$*Exposure-Foreign-Liab$_{f,pre}$	3.7067***				-0.9834			
	(1.234)				(1.008)			
Crisis$_{yq}$*Exposure-Local-Liab$_{f,pre}$	-0.4594				0.9091			
	(1.839)				(1.153)			
Post$_{yq}$*Exposure-Foreign-Liab$_{f,pre}$	-0.1352				-1.8852**			
	(1.225)				(0.785)			
Post$_{yq}$*Exposure-Local-Liab$_{f,pre}$	-0.7201				0.3475			
	(1.792)				(0.915)			
Crisis$_{yq}$*Exposure-Foreign$_{f,2007:Q1}$		2.3085*				0.1307		
		(1.292)				(0.911)		
Crisis$_{yq}$*Exposure-Local$_{f,2007:Q1}$		-1.6043				3.2150*		
		(2.334)				(1.678)		
Post$_{yq}$*Exposure-Foreign$_{f,2007:Q1}$		1.4143				-0.8845		
		(1.391)				(0.722)		
Post$_{yq}$*Exposure-Local$_{f,2007:Q1}$		0.9935				0.6597		
		(2.217)				(1.266)		
Crisis$_{yq}$*Exposure-Foreign-Log$_{f,pre}$			0.0823***				-0.0217	
			(0.030)				(0.023)	
Crisis$_{yq}$*Exposure-Local-Log$_{f,pre}$			0.0092				-0.0011	
			(0.031)				(0.019)	
Post$_{yq}$*Exposure-Foreign-Log$_{f,pre}$			-0.0362				-0.0547**	
			(0.032)				(0.023)	
Post$_{yq}$*Exposure-Local-Log$_{f,pre}$			-0.0010				-0.0084	
			(0.027)				(0.018)	
Crisis$_{yq}$*Exposure-Foreign$_{f,2005}$				4.1977**				-0.6689
				(1.978)				(1.431)
Crisis$_{yq}$*Exposure-Local$_{f,2005}$				0.5461				1.7655
				(3.353)				(1.784)
Post$_{yq}$*Exposure-Foreign$_{f,2005}$				-0.6775				-2.6736**
				(1.794)				(1.228)
Post$_{yq}$*Exposure-Local$_{f,2005}$				-2.0408				1.4118
				(3.198)				(1.562)
N	15269	15269	15269	15269	25294	25294	25294	25294
R^2	0.8477	0.8476	0.8476	0.8476	0.8395	0.8395	0.8395	0.8395
Firm Controls	YES	YES	YES	YES	YES	YES	YES	YES
Bank Controls	YES	YES	YES	YES	YES	YES	YES	YES
Firm FE	YES	YES	YES	YES	YES	YES	YES	YES
Industry*Year-Quarter FE	YES	YES	YES	YES	YES	YES	YES	YES
Companies active in both	Excluded	Excluded	Excluded	Excluded	Excluded	Excluded	Excluded	Excluded

Panel E: Collapsed Pre/Policy/Crisis Time Dimension – Imports and Exports

	(1) $Exports_{f,yq}$	(2) $Imports_{f,yq}$
$Crisis_{yq}$*Exposure-$Foreign_{f,pre}$	5.5847***	-1.7361
	(1.919)	(2.136)
$Post_{yq}$*Exposure-$Foreign_{f,pre}$	-1.8223	-3.5427**
	(2.322)	(1.504)
$Crisis_{yq}$*Exposure-$Local_{f,pre}$	-1.4024	2.6373
	(3.219)	(1.785)
$Post_{yq}$*Exposure-$Local_{f,pre}$	-1.9005	0.8679
	(3.108)	(1.345)
N	2859	4735
R^2	0.9522	0.9485
Firm Controls*[Post; Crisis]	YES	YES
Bank Controls*[Post; Crisis]	YES	YES
Firm FE	YES	YES
Industry*Year-Quarter FE	YES	YES

Panel F: Different Definitions of Crisis and Policy Periods – Imports and Exports

	(1) $Exports_{f,yq}$	(2) $Imports_{f,yq}$	(3) $Exports_{f,yq}$	(4) $Imports_{f,yq}$
$Crisis_{yq}$*Exposure-$Foreign_{f,pre}$	5.3119**	-1.8240	4.3898**	-1.5481
	(2.083)	(1.942)	(2.112)	(1.892)
$Crisis_{yq}$*Exposure-$Local_{f,pre}$	-1.5080	2.6928	-1.1866	2.0264
	(3.865)	(2.196)	(3.732)	(2.048)
$Policy_{yq}$*Exposure-$Foreign_{f,pre}$	-0.9619	-3.1646**	-0.1561	-3.4418***
	(2.257)	(1.254)	(2.328)	(1.283)
$Policy_{yq}$*Exposure-$Local_{f,pre}$	-1.3959	0.9558	-1.5456	1.5529
	(3.341)	(1.640)	(3.465)	(1.691)
N	14312	23708	15269	25294
R^2	0.8485	0.8395	0.8476	0.8395
Firm Controls*[Post; Crisis]	YES	YES	YES	YES
Bank Controls*[Post; Crisis]	YES	YES	YES	YES
Firm FE	YES	YES	YES	YES
Industry*Year-quarter FE	YES	YES	YES	YES
2008:Q3	Excluded	Excluded	Policy	Policy

Panel G: Excluding companies in oil-related sector – Imports and Exports

	(1) $Exports_{f,yq}$	(2) $Imports_{f,yq}$	(3) $Exports_{f,yq}$	(4) $Imports_{f,yq}$
$Crisis_{yq}$*Exposure-$Foreign_{f,pre}$	5.1864***	-0.6222	2.4086*	-0.2356
	(1.831)	(1.561)	(1.262)	(1.088)
$Crisis_{yq}$*Exposure-$Local_{f,pre}$	-1.8740	2.9694	-1.9950	1.7760
	(3.489)	(2.042)	(2.443)	(1.571)
$Post_{yq}$* Exposure-$Foreign_{f,pre}$	-0.5144	-2.8107**	1.7117	-3.0807***
	(2.315)	(1.253)	(1.550)	(1.021)
$Post_{yq}$*Exposure-$Local_{f,pre}$	-1.3191	1.5712	-1.9078	0.5051
	(3.414)	(1.636)	(2.391)	(1.258)
N	14200	24072	23698	35542
R^2	0.8466	0.8424	0.8739	0.8545
Firm Controls*[Post; Crisis]	YES	YES	YES	YES
Bank Controls*[Post; Crisis]	YES	YES	YES	YES
Firm FE	YES	YES	YES	YES
Industry*Year-Quarter FE	YES	YES	YES	YES

Panel H: Control group: companies inactive in FX-debt market (unaffected by CC)

	(1)	(2)	(3)	(4)
	$Exports_{f,yq}$		$Imports_{f,yq}$	
$Crisis_{yq}$*Exposure-$Foreign_{f,pre}$	3.6200*		-2.2746	
	(1.952)		(2.197)	
$Crisis_{yq}$*-$\Delta Liabilities_{f,2007}^{predicted}$		3.2271*		-2.2771
		(1.775)		(2.027)
$Crisis_{yq}$*-$\Delta Liabilities_{f,2007}^{residual}$		0.0794		-0.2093
		(0.286)		(0.230)
$Post_{yq}$*Exposure-$Foreign_{f,pre}$	-0.0309		-5.1293***	
	(2.429)		(1.650)	
$Post_{yq}$*-$\Delta Liabilities_{f,2007}^{predicted}$		-1.1613		-4.4119***
		(2.010)		(1.450)
$Post_{yq}$*-$\Delta Liabilities_{f,2007}^{residual}$		-0.4094		-0.1692
		(0.257)		(0.262)
N	17274	17274	35187	35187
R^2	0.8367	0.8367	0.8145	0.8145
Firm Controls	YES	YES	YES	YES
Bank Controls	YES	YES	YES	YES
Firm FE	YES	YES	YES	YES
Industry*Year-Quarter FE	YES	YES	YES	YES
Sample of Companies	Foreign + Inactive	Foreign + Inactive	Foreign + Inactive	Foreign + Inactive

Panel I: Only companies constrained by capital controls

	(1)	(2)	(3)	(4)
	$Exports_{f,yq}$		$Imports_{f,yq}$	
$Crisis_{yq}$*Exposure-$Foreign_{f,pre}$	4.1191**		-1.1893	
	(1.872)		(2.202)	
$Crisis_{yq}$*-$\Delta Liabilities_{f,2007}^{predicted}$		3.3117**		-1.6573
		(1.618)		(2.085)
$Crisis_{yq}$*-$\Delta Liabilities_{f,2007}^{residual}$		-0.0726		-0.1470
		(0.270)		(0.240)
$Post_{yq}$* Exposure-$Foreign_{f,pre}$	0.9091		-2.9797*	
	(2.646)		(1.411)	
$Post_{yq}$*-$\Delta Liabilities_{f,2007}^{predicted}$		0.0689		-2.3853*
		(2.227)		(1.323)
$Post_{yq}$*-$\Delta Liabilities_{f,2007}^{residual}$		-0.4443		-0.0804
		(0.271)		(0.233)
N	3956	3861	5640	5453
R^2	0.8343	0.8347	0.8106	0.8105
Firm Controls*[Post; Crisis]	YES	YES	YES	YES
Bank Controls*[Post; Crisis]	YES	YES	YES	YES
Firm FE	YES	YES	YES	YES
Industry*Year-Quarter FE	YES	YES	YES	YES

This table shows the effect of capital controls on firm-level trade, depending on pre-policy exposure to foreign and local FX-debt markets, during the implementation of the policy and the following Crisis. In Panel A and B, respectively, we report progressively saturated versions of the model for exports and imports. In Panel C, we perform the Oster (2017)'s test on the coefficient $Crisis_{yq}$*Exposure-$Foreign_{f,pre}$ ($Post_{yq}$*Exposure-$Foreign_{f,pre}$) for exports (imports) regressions, based on the comparison of columns 1 and 4 of Panel A (B) – under the assumption that the maximum R^2 is equal to 1. In Panel D, we check the robustness of results to different definitions of the exposure variables. In Panel E, we collapse data by taking averages of firm-level dependent and independent variables over the periods: 2006:Q1-2007:Q1 (pre); 2007:Q2-2008:Q2 (policy); 2008:Q3-2008:Q4 (crisis). In Panel F, we either exclude observations for 2008:Q3 (columns 1 and 2) or relabel them as a year-quarter with CC in place (i.e. with $Post_{yq}$ equal to 1 and $Crisis_{yq}$ equal to 1 in columns 3 and 4). In Panel G, we repeat baseline regressions excluding companies in involved in the production, distribution and refinement of oil (ISIC sectors 10, 11, 12, 13, 14, 23 and industries 2521, 2529 and 2924). In Panel H, we replicate regressions in Table 8, Panels A and B, contrasting the firm-level exports and imports of firms exposed to CC (i.e. firms ex-ante borrowing in FX from foreign banks only, whose growth of total liabilities is limited by the policy) and of firms inactive in the FX-debt market (unaffected by CC). In Panel I, we replicate regressions in Table 8, Panels A and B, based only on the sample of companies exposed to capital controls. List of Variables. $Exports_{f,yq}$ is defined as the logarithm of (1+Exports of firm f in period yq), $Imports_{f,yq}$ is defined as the logarithm of (1+Imports of firm f in period yq). Exposure-$Foreign_{f,pre}$ is the average of FX-Foreign $Inflows_{f,yq}$ over the period from 2005:Q1 to 2007:Q1; Exposure-$Local_{f,pre}$ is the average of FX-Local $Inflows_{f,yq}$ over the period from 2005:Q1 to 2007:Q1. Exposure-Foreign-$Liab_{f,pre}$ and Exposure-Local-$Liab_{f,pre}$ are the average firm-level FX debt inflows from foreign and local banks in the period 2005:Q1-2007:Q1, rescaled by total liabilities. Exposure-$Foreign_{f,2007:Q1}$ and Exposure-$Local_{f,2007:Q1}$ are given by the 2007:Q1 firm-level values of foreign and local FX-debt inflows over total assets. Exposure-Foreign-$Log_{f,pre}$ and Exposure-Local-$Log_{f,pre}$ are the average firm-level log FX debt inflows from foreign and local banks in the period 2005:Q1- 2007:Q1. Exposure-$Foreign_{f,2005}$ and Exposure-$Local_{f,2005}$ represent the average firm-level FX-debt inflow (rescaled by total assets) from local and foreign banks over the period 2005:Q1 to 2005:Q4. -$\Delta_{1y}Liabilities_{f,2007}^{predicted}$ is the yearly reduction in total liabilities predicted by Exposure-$Foreign_{f,pre}$ in a cross-sectional regression in 2007 with industry fixed effects. The residual heterogeneity in total liabilities from same regression is -$\Delta_{1y}Liabilities_{f,2007}^{residual}$. $Post_{yq}$ is a dummy with value 1 from 2007:Q2 onwards. $Crisis_{yq}$ is a dummy with value 1 from 2008:Q3 to 2009:Q4 and 0 from 2006:Q1 to 2008:Q2. Firm Controls include $ROA_{f,y-1}$, $Size_{f,y-1}$ and $Imports_{f,yq-1}$ ($Exports_{f,yq-}$) in regressions where exports (imports) is the dependent variable. Bank Controls include: $BankCET1_{f,yq-1}$; $BankROA_{f,yq-1}$; $BankSIZE_{f,yq-1}$; $BankNPL_{f,yq-1}$; $BankSaving_{f,yq-1}$; $BankChecking_{f,yq-1}$; $BankFX\text{-}Funds_{f,yq-1}$. Both Bank and Firm controls are fully interacted with the $Post_{yq}$ and $Crisis_{yq}$ dummies. Standard errors in parentheses are double-clustered at the firm and industry*year-quarter level. *** p<0.01, ** p<0.05, *p<0.1.

TABLE A9 – The Impact of Capital Controls on Industrial Employment

	(1)	(2)	(3)	(4)	(5)	(6)
			Employment$_{i,yq}$			
Crisis$_{yq}$ * Exposure-Foreign$_{i,pre}$	2.2977	3.4812	2.6434*	2.3676*	1.3606*	0.0357***
	(7.036)	(7.372)	(1.341)	(1.385)	(0.731)	(0.010)
Crisis$_{yq}$ * Exposure-Local$_{i,pre}$		0.0300	0.7628	1.1700	0.4724	0.0180
		(15.589)	(1.717)	(1.777)	(0.593)	(0.012)
Post$_{yq}$ * Exposure-Foreign$_{i,pre}$	-1.5892	-0.1510	-0.3714	-0.2195	0.1173	-0.0113
	(6.489)	(6.125)	(1.438)	(1.517)	(0.888)	(0.011)
Post$_{yq}$ * Exposure-Local$_{i,pre}$		1.9501	2.5174	2.8305	1.1627*	0.0230
		(15.246)	(2.100)	(2.117)	(0.686)	(0.015)
Exposure-Foreign$_{i,pre}$	-1.4168	-0.5089	-	-	-	-
	(4.466)	(3.999)				
Exposure-Local$_{i,pre}$		5.0855	-	-	-	-
		(10.128)				
N	432	432	432	432	432	432
R^2	0.0076	0.1777	0.9705	0.9732	0.9733	0.9754
Firm Controls*[Post; Crisis]	NO	YES	YES	YES	YES	YES
Industry FE	NO	NO	YES	YES	YES	YES
Time FE	NO	NO	NO	YES	YES	YES
Expo. Rescaling	Assets	Assets	Assets	Assets	Liabilities	Logs

This table shows the impact of capital controls on industrial employment. The dependent variable is defined as the logarithm of Employment in industry i in year-quarter yq. Exposure-Foreign$_{i,pre}$ is a proxy of industry-level exposure to foreign banks. In columns (1) to (4), this is computed as the weighted average of the mean FX-debt flow from 2005:Q1 to 2007:Q1 across firms; weights are given by the ratio between a firm total assets and total assets at the end of 2006. In column (5), FX-debt flows at the firm level are rescaled by total liabilities. In column (6), they are defined in logs. Similar measures are used for FX-debt flows from local banks, whose exposure is denoted by Exposure-Local$_{i,pre}$. Post$_{yq}$ is a dummy with value 1 from 2007:Q2 onwards and 0 from 2006:Q1 to 2007:Q1. Crisis$_{yq}$ is a dummy with value 1 from 2008:Q3 to 2009:Q4 and 0 from 2006:Q1 to 2008:Q2. Controls include ROA$_{i,y-1}$, Size$_{i,y-1}$, Exports$_{i,yq-1}$, Imports$_{i,yq-1}$. All controls are interacted with the Post$_{yq}$ and Crisis$_{yq}$ dummies. Standard errors in parentheses are clustered at the industry*Period level. Period is a categorical variable with value: 1 from 2006:Q1 to 2007:Q1; 2 from 2007:Q2 to 2008:Q2; 3 from 2008:Q3 to 2009:Q4. *** p<0.01, ** p<0.05, *p<0.01.

3 CAPITAL CONTROLS, DOMESTIC MACROPRUDENTIAL POLICY AND THE BANK LENDING CHANNEL OF MONETARY POLICY

Joint with Martha López Piñeros, José Luis Peydró and Paul Eduardo Soto

3.1 Introduction

Credit booms greatly amplify business cycle fluctuations and are the main predictors of financial crises, especially credit booms that are financed with foreign liquidity (Gourinchas and Obstfeld, 2012; Jordà, Schularick, and Taylor, 2011; Mendoza and Terrones, 2008; Reinhart and Reinhart, 2008; Schularick and Taylor, 2012). Macroprudential policies, including capital controls (CC), try to tame excessive credit booms. Since the Global Financial Crisis (GFC) of 2008-2009, macroprudential policies have become increasingly popular among both academics and policymakers (Freixas, Laeven and Peydró, 2015) and their use has risen constantly (Claessens, 2015; Alam et al., 2019). Moreover, the International Monetary Fund (IMF) has endorsed capital controls as a temporary and last resort tool for managing credit booms led by large capital inflows, especially when room for standard macroeconomic policy is exhausted (Blanchard, 2013; IMF, 2012, 2018; Ostry et al., 2010; Qureshi et al., 2011).

In the same spirit, a class of models rationalizes capital controls as a Pigouvian tax to reduce the negative externalities on systemic risk and aggregate demand associated with excessive foreign debt (Benigno et al., 2016; Bianchi, 2011; Brunnermeier and Sannikov, 2015; Jeanne and Korinek, 2010; Korinek, 2011, 2018; Korinek and Sandri, 2016; Schmitt-Grohé and Uribe, 2016). Other authors support capital controls based on the idea that controls insulate local monetary policy from shocks originated in global financial centers (Rey, 2015; Farhi and Werning, 2012, 2014, and 2016; Davis and Presno, 2017).

We analyze the effects of capital controls and domestic macroprudential policy on credit supply. For identification, we exploit the simultaneous introduction of capital controls on foreign exchange (FX) debt inflows and an increase of reserve requirements on domestic bank deposits in Colombia during a strong credit boom, as well as administrative credit registry and supervisory bank balance sheet data. In brief, we find the following robust results.

First, banks use cheaper FX-funding from abroad to arbitrage contractionary local monetary (interest rate) policy. An increase in the local monetary policy rate raises the interest rate differential with respect to the United States, allowing more FX-indebted banks to carry trade cheap FX-funds with expensive local credit supply. The carry trade is stronger during periods of relatively larger deviations from the

Covered Interest Parity (CIP) and amplifies bank risk-taking in lending, as it directs the supply of credit toward ex-ante riskier and relatively opaque local firms. Capital controls, by taxing FX-debt, reduce the interest rate differential and break the carry trade, enhancing the bank-lending channel of local monetary policy rates and reducing bank risk-taking.

Second, the increase in reserve requirements on domestic deposits directly reduces credit supply during the boom, and more so for riskier firms, rather than indirectly enhancing the effects of monetary rates on credit supply. Importantly, banks' reliance on domestic deposits and FX-financing are strongly negatively correlated, suggesting that those banks which restrict credit supply more due to capital controls are less influenced by the domestic reserve requirements, and the other way around. This implies that the two policies affect credit supply independently of each other and that both contribute to slowing down the credit boom.

Our main contribution to the literature is to show that both capital controls and domestic macroprudential policy tame credit supply booms, including credit supply to ex-ante riskier firms, by targeting different but complementary sources of bank debt. Capital controls target bank foreign funding, thereby improving the effectiveness of the bank lending channel of local monetary policy. Domestic macroprudential policy targets bank domestic debt, directly attenuating credit supply booms. As credit booms stem from *both* foreign *and* local liquidity, and we find that banks which finance the credit boom with domestic deposits *rely less* on foreign (FX) debt (and vice versa), our results suggest that a Tinbergen rule with two (macroprudential) instruments is necessary to tackle the two (intermediate) objectives (sources of liquidity). In other terms, the two macroprudential instruments target the two sources of bank debt, foreign and domestic, that drive the credit boom.

The remaining of this Introduction provides a detailed preview of the paper and a discussion of the related literature and its contrast with this paper.

DETAILED PREVIEW OF THE PAPER

We analyze two related research questions. First, we ask whether (and if so, why) capital controls (CC) on FX-financing strengthen the bank-lending channel of monetary policy by increasing the pass-through of variations in the *local* policy rate to domestic credit, and the implications of this for bank risk-taking. Second, we investigate the impact of domestic macroprudential measures on credit supply, in particular reserve requirements (RR) levied on bank local financing through household and firm deposits, as well as whether RR affects the impact of monetary policy rates on credit supply. By doing so, we can analyze whether the two macroprudential measures operate through different channels targeting respectively bank foreign or domestic liabilities, and whether they help mitigate credit supply booms, including the risks stemming from credit expansion during a boom.

Our work is based on two administrative datasets provided by the Colombian Financial Supervisory Authority. First, we have access to the National Credit Registry (CR), which collects detailed quarterly corporate loan information at the loan-level. The CR tracks information on the universe of commercial loans provided to nonfinancial companies. Second, we have access to bank supervisory quarterly balance sheets, which include data on bank size, profitability, capital, nonperforming loans (NPL), and, most importantly for our purposes, the volume of the sources of bank financing taxed through RR and CC (domestic deposits as well as FX funding).

For capital controls, we exploit the Central Bank of Colombia's introduction in May 2007, during a strong credit boom, of a 40 percent *unremunerated* reserve requirement (URR) on FX debt inflows. At the time, local interest rates – as reflected by the overnight interbank rate – were as high as 8.4 percent. Hence, the new regulation resulted in high taxation of FX debt inflows as a large part (40 percent) of the inflows were in the central bank as unremunerated reserves. CCs were borne by the banks to the extent that FX funding was raised to finance peso investments,[66] including lending, and were deposited for six months at the central bank without any remuneration; the deposit could be withdrawn before this deadline, but upon the payment of a heavy fee (decreasing in time and ranging from 9.4 percent of the deposit in the first month to 1.6 percent during the sixth and last month). CCs were lifted by October 2008 amid signs of an economic slowdown related to the unfolding of the GFC after Lehman Brothers' collapse.[67]

Concerning the domestic macroprudential measures, we exploit a contemporaneous policy change to traditional reserve requirements (RR) on peso-denominated deposits. In May 2007, the Central Bank introduced a *marginal* RR on bank deposits, on top of the ordinary reserve requirement, applied to the overall volume of new deposits received after May 7th, 2007. The marginal RR was not remunerated (at a time of high local interest rates) and was initially fixed at 27 percent for checking deposits and 12 percent for savings deposits, though it was eventually uniformed at 27 percent for both savings and checking deposits by June 2007.[68]

Both CC and RR are non-random, but rather induced by the credit boom, which affects both the demand and the supply of credit, i.e. both firms' and banks' financing and lending strategies. In this respect, we identify credit supply channels by exploiting variation in loan conditions for the same firm, in a given year:quarter, across banks with different exposure to either CC or RR. Put differently, we exploit ex-

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[66] When a bank's FX funding finances FX loans to local firms, the bank's customer pays the CC (in other terms, to avoid double taxation of capital inflows, bank FX funding is exempted). We also analyze FX loans to firms.
[67] Together with CC, the Central Bank fixed a cap on banks' gross FX-position (i.e. the sum of on- and off-balance-sheet FX assets and liabilities), equal to 500% of banks regulatory capital, which further constrained banks' ability to access FX-financing.
[68] At the time of the introduction of the *marginal* RR (May 2007), the level of the *ordinary* RR was 12% and 6% for checking and savings deposits, respectively, but it was eventually levelled at 8.3% in June of 2007.

ante heterogeneity in bank foreign (FX) funding and local funding (domestic deposits), respectively, as capital structures tend to be sticky over time. Therefore, we run loan-level regressions saturated with firm*year:quarter fixed effects, controlling for all idiosyncratic, observed and unobserved, time-varying unobserved shocks at the firm level (Khwaja and Mian, 2008).[69] Moreover, to understand the interaction of CC and RR with the local monetary policy rate, we further interact banks' exposures with local policy rates (Kashyap and Stein, 2000; Jiménez et al., 2012, 2014), as well as with the spread between the local and the US monetary rates. We also analyze the impact of macroprudential policies on risk-taking in credit supply (Jiménez et al., 2017).

Our main findings are as follows. We first evaluate the pass-through of the *local* policy rate variation on domestic bank credit over the three-year period from 2005Q2 to 2008Q2. The period ends before the Global Financial Crisis (GFC). In loan level regressions exploiting time variation only, we find that before the introduction of the macroprudential measures, an increase in the local policy rate is associated with positive subsequent growth in the volume of bank credit. However, after (compared to before) the implementation of the macroprudential policies, higher monetary policy rates imply lower credit volume.[70]

We next investigate a mechanism for explaining these findings. In principle, both CC and RR might influence the relation between the central bank's monetary policy rate and bank credit. Both local and foreign bank financing are in fact more expensive with the measures in place, which might render the influence of the local policy rate on bank credit more negative.

Our results show, however, that bank (foreign) FX-funding, not domestic deposit funding, drives the results. In particular, using the local-*versus*-U.S. policy rate differential, we find that before CC, banks with higher (versus lower) ex-ante FX funding increase their credit supply relatively more when the differential monetary rate goes up.[71] Moreover, after CC, these banks cut credit supply more sharply in reaction to an increase in the monetary interest rate differential. The effects are both statistically and economically significant. Before CC, following a 1 percentage point (p.p.) increase in the monetary policy interest rate spread, banks with a 1 standard deviation (s.d.) higher FX funding increase lending (to the same firm in the same quarter) by 3.8 p.p.. After CC, however, the same variations in the monetary rate spread and FX funding are associated with a relative reduction in credit by 3.5 p.p. These findings

[69] Alternatively, in robustness checks, we control for borrower demand via industry*time fixed effects for the sample of all firms, so to include as well those firms indebted with only one bank, which are excluded from the application of firm*time fixed effects.
[70] Overall, a higher monetary policy rate after the introduction of both macroprudential policies is associated with at most a non-positive reaction of bank credit, consistent with the strong credit boom that Colombia was experiencing.
[71] We do not find evidence of a significant interaction of exposure to RR with the local interest rate policy, or the difference between the local and the US monetary rates.

are consistent with a carry trade strategy by local banks, which borrow cheaply in FX to lend at higher rates in pesos, and with CC breaking such carry by strongly increasing the cost of bank FX borrowing.[72]

The carry trade affects local companies heterogeneously. We sort firms according to several proxies of *pre-policy* riskiness: the average interest payments on bank loans, the average share of bank loans with short maturity, i.e. below one year, a dummy variable for whether a firm ever defaulted on a bank loan, and a dummy variable describing whether a company's balance sheet is publicly supervised or not, which we interpret as a proxy for firm's opaqueness. Consistently across the different risk measures, we find that the pre-CC expansion in credit supply due to carry trade favors relatively riskier and opaque firms, whose credit also suffers a sharper reduction after the enforcement of CC.

To further understand the mechanism behind our results we provide two additional tests. First, we rerun the analysis over a representative subsection of credit registry loans for which we can access the breakdown by currency (peso versus FX). We find that the just-described results are driven by corporate lending in peso. This finding is reassuring for two reasons: i) reinvesting globally borrowed FX funds in local peso loans rather than FX loans grants higher returns, given the positive policy risk-free rate differential; ii) banks would bear the CC tax only if FX funds were reinvested in peso-denominated assets, so that credit supply variations induced by CC must show up among peso loans.[73] As a second test, we substitute the Colombia-U.S. policy rate spread with the deviations from the CIP computed by Du and Schregher (2016) over the three-month sovereign yield spread between Colombia and U.S. rates. Indeed, our results go through, i.e. positive variations in deviations from the CIP are associated with a relative jump (descent) in credit supply by higher FX-indebted banks before (after) the introduction of CC. This finding is important, as banks mostly hedge their FX liabilities and CIP deviations grant carry trade returns on top of the costs associated with hedging.[74]

[72] While the main focus of our paper rests on the interaction of CC with local interest rate policy, we also provide evidence that CC halt the dependence of domestic bank credit from global shocks. We show that absent CC, a tightening of global liquidity conditions (as proxied by a jump in the VIX), and/or a fall in global demand (captured by a decline in oil prices), triggering a depreciation of the Colombian currency, pushes more FX-indebted banks to cut credit. The introduction of CC reduce those effects, therefore dampening the implications of global shocks for bank credit. These findings align, among others, to the cross-country aggregate evidence in Zeev (2017), who shows that output is less sensitive to global credit supply shocks in countries with CC in place. For cross-country aggregate evidence against this hypothesis, see Bergant et al. (2020).

[73] Note that if banks borrowed in FX to finance FX loans, the CC would be borne by the ultimate borrower, i.e. a local company. The reduction in FX loans associated with CC would show up as a demand shock, which is controlled for in our empirical setting by firm*year:quarter fixed effects (Khwaja and Mian, 2008).

[74] The effect of CIP deviations also survives if we additionally include the interaction of bank FX funding with the component of the three-month sovereign spread which is not accounted for by CIP-deviations, i.e. the three-month forward premium. Note that in our sample period CIP deviations are relatively small as compared to those observed after the Global Financial Crisis, but nonetheless account for roughly 17 percent of the mean sovereign yield spread between Colombia and the United States.

Our estimates suggest that carry trade lending implies higher bank risk-taking. During the boom, it increases the leverage of risky and opaque companies, which are likely to suffer more during a subsequent bust. At the same time, banks finance this risk-taking through FX non-core liabilities, which tend to be more fragile (Dagher and Kazimov, 2015; Demirgüç-Kunt and Huizinga, 2010; Hahm, Shin and Shin, 2013; IMF, 2019c; Ivashina, Scharfstein and Stein, 2015). In this respect, CC reduce banks' risk-taking (on assets and liabilities), on top of enhancing the bank-lending channel by halting one way for arbitraging local monetary rate policy.

After the introduction of RR, banks with higher ex-ante exposure to RR (higher ex-ante reliance on savings deposits and checking deposits) cut credit supply – i.e. reduce lending to the same firm at the same time compared with banks less exposed to RR. Moreover, this reduction in credit supply is robust to controlling for exposure to CC, or to CC and its interaction with the local policy rate, a first suggestion that the two macroprudential channels affect credit supply through two distinct channels. The robust results suggest a large economic impact of RR on bank credit supply. A 1 s.d. overall increase in ex-ante deposits affected by the RR shock (i.e. the sum of checking deposits and saving deposits) implies a 5.4 p.p. reduction in bank credit supply. Moreover, the RR policy change exerts heterogeneous effects across firms, with riskier and more opaque companies significantly more affected.

A final question is whether CC and RR affect different lenders through distinct channels, and whether both instruments are necessary to tame credit booms and associated risk-taking. The scatterplot in Figure 1, reporting bank reliance on local savings deposits and checking deposits on the x-axis and bank FX funds on the y-axis (both measures are expressed as a share of total assets), indicates that this is the case. Banks more exposed to RR (domestic bank deposits) are *less* exposed to CC (FX bank funding), and over the period of analysis, the two variables are correlated negatively (by a factor of 37%, significant at a 1% level).[75] Hence, CC and RR – i.e. macroprudential measures targeting foreign and domestic bank debt, respectively – affect credit supply through different channels.

RELATED LITERATURE

We contribute to several strands of literature. First, as we show that capital controls, by reducing banks' carry trades, increase the effectiveness of variations of the *local* monetary policy rate on bank credit supply, we contribute to the large literature on the bank lending channel (e.g. Bernanke and Gertler, 1995; Kashyap and Stein, 2000; Jiménez et al., 2012 and 2014; Acharya et al., 2020), including the related literature on international finance and monetary policy (e.g. Bräuning and Ivashina, 2020, and

[75] The two channels operate independently and are both significant in regressions in which both channels are allowed to affect bank credit.

forthcoming; Bruno and Shin, 2015a, 2015b; Cetorelli and Goldberg, 2012; Morais et al., 2019; Rey, 2015).

Several studies investigate empirically the extent of monetary policy autonomy depending on the degree of capital account openness, often in a cross-country framework (e.g. Klein and Shambaugh, 2015; Han and Wei, 2018). We contribute to this literature by showing a specific mechanism through which capital inflows reduce the pass-through of *local* interest rate policy to domestic credit, namely carry trade strategies by domestic banks.[76] This finding is consistent with recent theoretical insights from Cavallino and Sandri (2019) that a local monetary contraction – by widening the interest rate differential between a small open economy and the rest of the world – drives carry trade inflows, which can significantly increase credit and risk in the local economy. On the empirical front, Fendoglu, Gulsen and Peydró (2019) document, in accordance with our results, that carry trade inflows on the interbank market impaired the bank lending channel of monetary policy in Turkey. Crucially, we show that capital controls are effective in breaking the carry trade, thereby contributing to increase the pass-through (effectiveness) of domestic monetary policy rates to bank credit supply (postulated by Rey, 2015). Furthermore, we find that CC reduce bank risk-taking in both bank assets and liabilities. The pre-CC bank carry trade, driven by the local interest rate policy, increases credit supply to the ex-ante riskier and more opaque firms, and banks finance this risk-taking with FX fragile funding. This result depicts a previously overlooked but nonetheless important prudential mechanism of CC, especially beneficial in light of the poor performance of carries during major financial downturns, including the GFC (Koijen et al., 2018).

Closer to our paper, Dias et al. (2021) exploit the Colombian CC in 2007 to analyze the relation between capital controls and monetary policy. Similarly to us, they conclude that CC strengthen the transmission of monetary policy rates on lending. However, our focus is different, centered on the influence of *local* (as opposed to *international*) monetary policy. We also analyze a particular mechanism, namely banks' carry trade from cheaper FX funds to the supply of credit in higher-rate peso loans (and even more to riskier and opaque local firms), which we show to be especially reactive to the difference between *local* and *international* policy rates. Moreover, we analyze the interaction with different macroprudential policies, finding that capital controls and domestic macroprudential policy complementarily mitigate the boom and the associated risk-taking through two distinct channels, independently operating through global and domestic liquidity, respectively.

[76] Other studies focus on carry trades by large nonfinancial companies (NFCs) in Emerging Markets (Acharya and Vij, 2016; Caballero, Panizza, and Powell, 2016; Bruno and Shin, 2017), and highlight how their U.S. dollar debt increases when carry trade is more favorable. Liao (2020) shows that carry trade explains a large fraction of international bond issuance. Differently, our attention rests on carry trades by domestic banks in Emerging Markets, involving local currency loans to domestic NFCs, including SMEs.

Our paper additionally speaks to a growing literature on the deviations from CIP.[77] In particular, consistent with our findings, Avdjiev et al. (2019) document that a stronger USD is associated with significant CIP-deviations and with a reduction in USD-denominated cross-border banking flows. We show that CIP-deviations can hamper the transmission of local policy rate hikes to domestic credit and that CC are eventually useful to enhance such transmission.

We also contribute to the literature by showing complementarities between domestic macroprudential policies and capital controls, highlighted theoretically by Korinek and Sandri (2018). Credit booms stem from both local and foreign sources of liquidity, with the latter flowing to the local economy either through foreign lending or through domestic bank international non-core FX funding (Avdjiev, McCauley and McGuire, 2012; Borio, McCauley and McGuire, 2011; Hahm, Shin and Shin, 2013). We show that CC tame credit booms because, by targeting foreign bank debt, they increase the effectiveness of domestic interest rate policy on credit supply. However, CC do not target domestic liquidity -e.g. bank deposits from local households and firms- that constitute the bulk of domestic bank funding. We show that domestic macroprudential policy via (tightening of) RR cuts credit supply by targeting domestic bank deposits. The increase in RR on domestic deposits directly reduces credit supply during the boom, and more so for riskier firms, rather than (indirectly) enhancing the effects of local monetary rates on credit supply.[78]

Overall, our results innovate the literature on macroprudential policy (see e.g. Galati and Moessner, 2013; 2018) by suggesting a "prudential Tinbergen rule" for tackling booms driven by a combination of domestic and foreign liquidity that is used by different financial intermediaries for financing their lending activities. Two instruments, i.e. CC and one domestic prudential measure –RR in our Colombian episode– are necessary to tackle the two (intermediate) objectives (sources of liquidity).

The rest of the paper is organized as follows. Section 2 describes the two policy changes and the datasets. Section 3 presents the results on the bank lending channel of monetary policy. In Section 4, we discuss findings on domestic reserve requirements. Section 5 briefly concludes.

[77] For evidence on deviations from CIP in both Advanced and Emerging Economies, see e.g. Borio, McCauley and McGuire, 2016; Cerutti, Obstfeld and Zhou, 2019; Du and Schreger, 2016; Du, Tepper and Verdelhan, 2018.
[78] We analyze CC in conjunction with other domestic RR-policies, highlighting different channels of transmissions to credit supply, whereas most existing studies focus on just one of the two policies. For evidence on the effectiveness of prudential RR, see, among others, Barroso et al. (2020), Cordella et al. (2014) and Federico, Vegh and Vuletin (2014).

3.2 Institutional Settings and Data

3.2.1 Capital Controls on Capital Inflows and Reserve Requirements Policy in Colombia

The Colombian economy expanded rapidly in the mid-2000s, with annual GDP growth above 4 percent in both 2004 and 2005. At least from early 2006, inflationary pressures further intensified due to a pronounced surge in domestic credit. The annual growth rate of commercial credit more than doubled in 2006- from less than 10 percent to 22 percent (Figure 2, Panel A). The Central Bank reacted by steadily increasing the interest rate, from 6 percent at the end of 2005 to 8 percent by early 2007 and further up to 10 percent in mid-2008. A higher monetary policy rate was accompanied by a widening interest rate differential *vis-à-vis* the U.S. Fed Funds Rate as early as mid-2006 (Figure 2, Panel B). These developments triggered strong capital inflows - especially non-FDI debt inflows - by third quarter 2006 (and peaking in first quarter 2007 just before the introduction of capital controls), as well as an associated sharp appreciation of the Peso-USD nominal exchange rate.

To deal with the acceleration of domestic credit boom, financed in part with foreign liquidity, the Central Bank resorted to a package of unconventional prudential measures on May 7th, 2007.

First, Capital Controls were introduced in the form of an Unremunerated Reserve Requirement (URR) on all new FX debt inflows.[79] The URR works as follows: upon disbursement of the FX credit to a Colombian firm (either a bank or a nonfinancial company), 40 percent of the nominal loan amount is deposited in an account at the Central Bank, with no remuneration in return. The deposit is always borne by the ultimate borrower of the debt and can be withdrawn without penalty only after six months. At the time, local interest rates –as reflected by the overnight interbank rate– were as high as 8.4 percent. The new regulation resulted in high taxation of FX debt inflows.[80] CC were borne by the banks to the extent that FX funding was raised to finance peso investments, including lending. When bank FX funding finances FX lending, it is the bank's customer that pays the CC (to avoid double taxation of capital inflows, bank FX funding is exempted).[81] In this paper, we focus on the impact of CC on domestic credit through a bank-financing channel, where most firms in Colombia are small and medium sized enterprises (SMEs) without access to FX corporate debt.[82] Finally, CC were lifted by October 2008 amid

[79] Portfolio inflows were initially excluded, but eventually made subject to the URR just one week after. On the contrary, foreign direct investments (FDI) were not subject to the URR, though in May 2008 a minimum stay of 2 years was applied to FDI.

[80] Earlier withdrawals were allowed with the payment of a heavy penalty. The penalty decreased in time and ranged from 9.4 percent of the deposit in the first month to 1.6 percent during the sixth and last month.

[81] Colombian banks and banks from other countries that follow Basel capital rules basically fully hedge their FX exposure.

[82] In the second chapter of this thesis, we analyze the effects of CC directly borne by non-financial companies, focusing in particular on the subsample of roughly 1,200 (large and export-oriented) firms issuing FX-debt without

signs of an economic slowdown related to the unfolding of the GFC after Lehman's collapse. Moreover, joint with CC, the Central Bank introduced an upper bound on the banks' gross FX-position (i.e. the sum of on- and off-balance-sheet FX assets and liabilities), equal to 500% of banks regulatory capital. This constrained further banks' ability to access FX-financing.

Contemporaneously with the CC, the Central Bank also modified its policy on Reserves Requirements (RR) on bank domestic financing. In May 2007 the Central Bank introduced a *marginal* RR on bank deposits, to be applied on top of the ordinary reserve requirements to new deposits received after May 7th, 2007. In other terms, the *marginal* RR would only apply on the increase in total bank deposits after May 7th, 2007. The marginal RR was not remunerated (at a time of high local interest rates) and was fixed at 27 percent for checking deposits and 12 percent for savings deposits. At the time of the introduction of the *marginal* RR, the level of ordinary RR was 12 percent and 6 percent for checking and savings deposits, but it was eventually raised to 8.3 percent just one month later in June 2007 – contemporaneously, the marginal RR was set at 27 percent for both savings and checking deposits. The marginal RR was eliminated in August 2008.[83]

3.2.2 Data and Summary Statistics

Our work is based on two administrative datasets provided by the Colombian Financial Supervisory Authority (*Superintendencia Financiera de Colombia*). First, we have access to the National Credit Registry (CR), which collects detailed quarterly information at the loan level for corporate loans, with information on loan volume and other loan characteristics.[84] The CR tracks information on the universe of commercial loans provided to nonfinancial companies. We aggregate loan-level data at the firm*bank level, by computing the total debt provided by a given bank to a company in a given year:quarter. Second, we have access to bank supervisory quarterly balance sheets, which include data on bank size, profitability, capitalization, nonperforming loans (NPL), and, most importantly for our purposes, the volume of the sources of bank financing taxed through RR and CC, i.e. domestic deposits and foreign FX inflows, respectively. The two datasets are matched through unique banking group identifiers.

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credit intermediation by banks operating in Colombia. For comparison, the largest sample in this paper comprehends 110,226 companies.

[83] In 2007, regulators also introduced changes with respect to loan provisions. Countercyclical loan provisions were introduced in July 2007 and the criteria was that each financial institution must accumulate or deplete its countercyclical provisions according to four criteria: deterioration of portfolio, efficiency, fragility and loan growth. In addition, in May 2007, there was a change in the rule for computing banks' loan losses provisions, based on expected rather than incurred losses. Throughout the paper, we show that our findings are not significantly affected by such policy change (whose effects are investigated by López, Tenjo and Zárate, 2014, and Morais et al., 2020).

[84] For each loan, we observe the interest payments (not interest rates) to proxy for credit risk and an indicator for whether the maturity of the loan is less than one year to proxy for liquidity risk.

We report the summary statistics in Table 1. In Panel A, we show the summary statistics for the largest sample we analyze throughout the paper, referring to regressions where we exploit time-variation to measure the unconditional impact of local monetary policy rate on bank credit. In this setting, we apply at most firm*bank fixed effects and bank controls. Therefore, the only requisite for a firm*bank pair to enter the sample is that it appears twice in the CR during the period of analysis 2005Q2-2008Q2.[85] This leaves us with 110,226 companies and 40 banks, corresponding to 12 major banking groups. Throughout the different year:quarters, this sample accounts on average for about 90 percent of total commercial credit.[86] $Loan_{f,b,yq}$ is expressed as the log of total outstanding (end of quarter) firm-(f)*bank-(b) debt, expressed in Colombian pesos as of 2005Q1. To get a sense of the magnitude of loans, the average loan is roughly 8,500USD as of 2005Q1. There are large differences in loan size across companies, though. A one interquartile variation in loan size reflects larger loans by more than 40,000USD as of 2005Q1.

Throughout our period of analysis, the monetary policy rate, labelled as $i_{yq-1,}$ is close to 7.5 percent on average. We also employ other measures of interest rate policy, including the growth of the local policy rate over a half year and over one year. Additionally, we use Taylor rate residuals, derived from two different rules: one expressing the policy rate as a function of the lagged yearly inflation rate and output gap (Rule 1), and the other as a function of yearly inflation and log GDP (Rule 2). A further important measure in our analysis is given by the spread between the local policy rate and the effective US FED Funds Rate, i.e. $MPspread^{US}_{yq-1.}$ Throughout the period of analysis, the spread is constantly positive and is about 3 percent on average. The distribution of the spread between the 3-month sovereign Colombian and U.S. yields mirrors very closely that of $MPspread^{US}_{yq-1}$ (augmented by a half p.p. premium). A factor explaining the sovereign spread may be deviations from the CIP. Although the largest deviations are observed after the Financial Crisis (see, e.g., Borio et al., 2016), they are still significant throughout our period of analysis and amount on average to 17 percent of the mean sovereign yield spread.[87] Finally,

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[85] Note: CC were removed in early October 2008, i.e. in 2008Q4. Nonetheless, we always stop our sample in 2008Q2 to avoid contaminating the effects of capital controls with those of the 2008-2009 Global Financial Crisis, which implied a sharp increase in the volatility of capital inflows (Forbes and Warnock, 2012) and which unfolded beyond the US borders after the failure of Lehman Brothers in mid-September 2008. All results on CC presented below are robust to the inclusion of observations for 2008Q3 in our samples. Moreover, our main results are even qualitatively robust (and, if anything, quantitatively stronger) after restricting the sample to 2007Q3, despite the significant reduction in the heterogeneity in monetary policy rates and banks' FX-financing (note that the Colombian central bank was raising the interest rate during this period). Additional cross-country (time-series) analysis based on BIS data shows that credit in Colombia slowed down significantly after 2007Q2, relatively to other Emerging Economies, including their subsample from Latin America. Results based on different samples (either shorter or longer) are available upon request.

[86] We exclude both financial companies and public utilities from the analysis, which roughly account for 10% of total commercial credit.

[87] The deviations from the CIP are retrieved from Du and Schreger (2016). In particular, they note that - absent CIP deviations - at a given tenor, the Colombia-US sovereign yield spread should equal the forward premium applied on a cross-currency swap that: i) buys U.S. zero-coupon Treasuries out of Colombian Peso; and ii) allows later on to enjoy the cash flows from US Treasuries in Colombian Pesos. Hence, they compute such forward premium (labelled as FP_{yq-1} in Panel B of Table 1) and subtract it from the Colombia-U.S. Sovereign yield spread, obtaining a series of deviations from the CIP. Du and Schreger (2016) compute those deviations for different

throughout all regressions on monetary policy rates, we apply further lagged macro controls, namely the annual growth rate of GDP, the lagged CPI index with base in 2005Q1 and the log Peso-USD exchange rate (expressed as Colombian Pesos per 1USD, so that an increase corresponds to a depreciation of the local currency).[88]

Panel B shows summary statistics for the smaller sample we focus on for the investigation of carry trade lending strategies triggered by variations in the local monetary policy rate. In this framework, we saturate the model with firm*year:quarter fixed effects, which excludes companies borrowing from one bank only in a given year:quarter, explaining the drop in observations with respect to Panel A. Note that companies with multiple lending relationships are typically larger, reflected by the fact that average loan size almost doubles. Indeed, the smaller sample of 37,867 multibank companies in Panel B represents a very large share of total commercial credit, close to 80 percent on average in our period (and, in turn, 90 percent of the aggregate credit in the sample of Panel A). Credit supply channels identified from regressions run over this sample therefore provide a representative picture of macroeconomic developments in bank credit. Regarding bank-level variables, the average FX indebtedness, denoted by $\text{FX-Funds}_{b,yq-1}$, equals 4.6 percent of total assets in the period from 2005Q2 to 2008Q2. This is a relatively large figure, larger for instance than the average common equity capital ($\text{CET}_{b,yq-1}$) over the same period, and more than half of the minimum threshold for regulatory capital (summing up Tier 1 and Tier 2 capital), fixed at 9 percent of total assets.

Importantly, the distribution of bank $\text{FX-Funds}_{b,yq-1}$ displays large heterogeneity, with a s.d. of 2.59 p.p. Nonetheless, the bulk of bank liabilities is given by domestic liquidity. In particular, savings deposits, denoted by the variable $\text{SavingD}_{b,yq-1}$, finance on average more than a third of a bank's total assets, whereas checking deposits (i.e. current accounts) -represented by the variable $\text{CheckingD}_{b,yq-1}$- fund 13.6 percent of total assets on average. Further, we have data on bank size (i.e. log total assets), nonperforming loans (i.e. loans at least 30 days past due, accounting on average for 2.7 percent of total loan volume at the bank level), and return on assets, which are quite homogenously distributed across banks and equal 1.4 percent on average on a quarterly basis.

Finally, for analyzing risk-taking associated with carry trade lending, we build various indicators of firm-level riskiness and opaqueness. First, we proxy credit risk through the average yield paid by a company over the pre-policy period 2005Q1 to 2007Q1, proxied through interest payments (rescaled by

sovereign bond tenors. We retain data for the 3-month tenor for two reasons. First, such data are available throughout the entire period of analysis for Colombia. Second, there is a tight link between the Colombian-US monetary policy rate spread and the 3-month sovereign yield spread. We aggregate data at the quarterly level by taking the average of the daily values.

[88] The Peso-USD exchange rate substantially correlates with both the VIX, reflecting the large influence of global liquidity conditions on the Colombian external sector, and with the oil price - which we alternatively use in some regressions – reflecting Colombia's dependence on oil exports.

loan size) and denoted by Firm $Risk_{f,pre}$. This is computed by taking, in each year:quarter, the weighted average of the loan-level "yields", with weights given by the loan shares relative to the total volume of bank debt at the firm level. Next, we take a firm-level average across the period 2005Q1 to 2007Q1. Note that an interquartile variation in such a variable corresponds to a 6.1 p.p. increase in the average firm-level yield, i.e. a 43 percent increase relative to the mean value, which we interpret as a sizable magnification of credit risk. The average reliance on short-term debt is computed with an analogous procedure, which is based on a 0/1 dummy for whether a loan has maturity no longer than one year. Companies rely on short-term debt for roughly one third of their total borrowing on average. The distribution, however, reveals significant differences across firms. A one interquartile variation implies higher reliance on short-term debt by a factor of 46.8 p.p. Firms in the fourth quartile of the distribution have more than half of their total debt with outstanding maturity below or equal to one year. These figures reflect large heterogeneity in refinancing risk across companies. As an additional measure for firm (default) risk, we also build a dummy with a value of 1 if a company has one or more loans with payments at least 30 days past due over the period 2005Q1 to 2007Q1, and 0 otherwise. In fact, the average value for this dummy shows that roughly 30 percent of the loans in our sample are granted to firms with such past due payments. Finally, a firm's opaqueness is proxied by a 0/1 dummy for whether a company's balance sheet is supervised by a public authority or not in the pre-policy period,[89] under the implicit assumption that balance sheet disclosure enhances firm transparency. Supervised companies represent about 10 percent of the firms in our sample, but they nonetheless account for about 30 percent of the loans, suggesting that those firms are larger and have more relationships in place with banks operating in Colombia.

Panel C reports the summary statistics for the sample we consider in the analysis of the RR policy. In this case, we run a traditional difference-in-differences exercise, comparing the evolution of bank credit before and after the introduction of the policy across differently exposed banks. Since shocks to the RR take place over the period 2007Q2-2008Q2, we build symmetric pre/post five quarter windows by running regressions over the year:quarters from 2006Q1 to 2008Q2. Again, as we isolate credit supply channels by saturating the model with firm*year:quarter fixed effects, the sample includes companies with at least two banking relationships in each year:quarter. To measure a bank's exposure to the RR shocks, we fix bank-level variables at their 2007Q1 value, the year:quarter preceding the shocks. We consider both savings and checking deposits alone, and their sum, denoted by the variable RR-$Depo_{b,2007Q1}$, which provides a measure of a bank's overall reliance on the liabilities targeted through the RR policy. The sum of checking and saving deposits accounts for nearly half of bank total assets in 2007Q1. In general, the distribution of all bank balance sheet items in 2007Q1 is very similar to that

[89] Companies with sufficiently large size, as measured by total assets, must disclose their balance sheet to Colombia's Authority for Supervision of Corporations (*Superintendencia de Sociedades*). Such data are also publicly available at the Authority's website.

described above for the longer period 2005Q2-2008Q2, suggesting a substantial stickiness in bank capital structure.

3.3 Capital Controls and the Bank-Lending Channel of Monetary Policy

In this section, we analyze how prudential measures affect the transmission of the local policy rate to bank credit. First, we verify that in the period of enforcement of CC and RR, the transmission is stronger, i.e. an increase in the local policy rate has a more negative effect on bank credit. Second, we ask whether the two policies are responsible for the enhancement of the bank-lending channel of monetary policy. Third, we investigate the implications of such lending strategy for bank risk-taking, and the eventual influence of CC on it.

3.3.1 Local Monetary Policy Rate and Bank Credit

We investigate the transmission of the local policy rate to bank credit through the lens of a loan-level regression model which exploits time variation over the period 2005Q2-2008Q2 within a given firm-bank pair. The most robust version of the model follows:

$$Y_{f,b,yq}=\beta_1 i_{yq-1} + \beta_2 Post_{yq} + \beta_3 Post_{yq}*i_{yq-1} + \gamma_1 MacroControls_{yq-1} + \gamma_2 BankControls_{b,yq-1} + \delta_{f,b} + \varepsilon_{f,b,yq}$$

The dependent variable, $Y_{f,b,yq}$, is the log total volume of outstanding debt provided by bank b to firm f in year:quarter yq (i.e. $Loan_{f,b,yq}$). The main coefficient of interest is β_3, describing the additional marginal effect of the lagged local policy rate i_{yq-1} on bank credit after the enforcement of CC and RR, on top of the pre-policy marginal effect, captured by β_1. $Post_{yq}$ is a dummy variable with value 1 from 2007Q2 onward and with value 0 before.

As the monetary policy rate is influenced by macroeconomic developments, which can affect bank credit as well, we include a vector of macro controls, $MacroControls_{yq-1}$, which may determine the Central Bank's policy reaction function. The Colombian monetary policy rate is formally governed by a pure inflation targeting regime so that we employ the lagged annual GDP growth rate and level of price, proxied by the CPI. Moreover, we add the lagged log exchange rate, controlling for the eventual influences of external factors (e.g. the dynamics of the Balance of Payments) on the local policy rate. The model is further augmented with a vector of lagged bank controls, consisting of bank FX funding, savings and checking deposits, size, ROA, common equity, and NPLs. We saturate the model with firm*bank fixed effects, denoted by $\delta_{f,b}$, which take care of all (observed and unobserved) time-invariant heterogeneity at the level of the single lending relationship. Finally, $\varepsilon_{f,b,yq}$ is an error term. We double-cluster standard errors at the firm and bank*(four-digit SIC)-industry level, a convention we maintain throughout the paper. Hence, we allow for correlation of the error-term both within-borrower (across time and lenders) *and* within-lender (across time and firms of a given industry).[90]

Table 2 shows the results from the estimation of the model. We report coefficients under progressively saturated versions of the model. In particular, in column 1, we employ just firm fixed effects, needed as a minimal set of controls to account for differences in the size of loans across firms. The related coefficients imply that, in the pre-policy period, a 1 p.p. increase in the local policy rate is associated with a jump in loan volume of 2.9 p.p. After CC and RR are enforced, however, the relation becomes negative,[91] which corresponds to an enhancement of the bank-lending channel. From a qualitative perspective, these relations are robust across all versions of the model and also survive the addition of firm*bank fixed effects, which increase the R-squared by 15 p.p. In the most saturated version of the model in column 5, corresponding to the equation commented above, before the introduction of CC and RR, a 1 p.p. increase in the local interest rate is associated with an expansion in loan volume of 3.5 p.p. After their introduction, however, we find again that the relation is more negative, and the interest rate does not overall exert a significant impact on loan volume. In other terms, irrespective of the model we consider, the results suggest that the introduction of the prudential measures contribute to strengthening the bank-lending channel of monetary (interest rate) policy.

In Table 3, we perform several robustness checks. In Panel A, we estimate alternative specifications of our model. In column 1, we modify the baseline model (in column 5 of Table 2) by further including firm*quarter(seasonal) fixed effects, which account for firm-specific seasonal demand shocks. In column 2, we control for the lagged loan-level provision for losses, rescaled by the loan amount. Contemporarily to the prudential shocks, a modification of the accounting rules for computing loan loss provisions was introduced and we aim to show that this does not interact significantly with our findings. In column 3, we rerun the baseline model weighting observation by the log loan size, so that the estimated coefficients do not reflect variations in very small loans. We augment the baseline weighted least squares model by progressively including again firm*quarter fixed effects in column 4 and loan loss provisions in column 5. Reassuringly, across all such model specifications, the period characterized by CC and RR is associated with a stronger negative effect of the local policy rate on bank credit, so that the qualitative interpretation of our findings does not change. We estimate the baseline model under

[90] We cluster at the bank*industry level rather than at the bank level because the latter option would leave us with less than 50 clusters, the conventional threshold for the minimum number of clusters (Cameron and Miller, 2015) which grant that asymptotic properties of the variance-covariance matrix estimator kick in. By taking the interaction of bank and (four-digit SIC) industry dummies, we obtain 4,246 clusters. Note that estimating standard errors interacting bank dummies with less granular (two- or three-digit SIC) industry dummies would leave the significance of our results unchanged. Still, we use four-digit SIC industry dummies as we also use such variables as fixed effects throughout the paper. Moreover, we show in robustness checks that our results survive under more conservative clustering strategies.

[91] We do not report the p-values of the test with null hypothesis: $\beta_1+\beta_3=0$. In columns 1 to 4 the p-value is constantly below 0.05, whereas in column 5 it is above 0.10.

alternative clustering strategies at the level of firm and bank in column 6, firm and bank and year:quarter in column 7, and firm and bank*industry and year:quarter in column 8, and find that our coefficients of interest are nonetheless significant at conventional levels.

Next, in Panel B, we verify that these findings hold across different proxies of the monetary policy rate. Importantly, in column 1, the increased pass-through to bank credit is robust to substituting the local policy rate with the spread between the local policy rate and the U.S. Effective Federal Funds Rates. This measure controls for an eventual dependence of local policy rates with cycles in U.S. monetary policy rates; the related coefficients imply that after the introduction of CC and RR, "purely" local positive interest rate variations reduce bank credit (by 0.57 p.p. in reaction to a 1 p.p. expansion in the spread). Generally, though, results are consistent across the different proxies of local interest rate policy. Note in particular that when employing Taylor residuals, we also obtain a full restoring of the bank-lending channel in the ex-post period; that is, before the introduction of CC and RR the local policy rate is positively linked to loan volume dynamics, and negatively thereafter (columns 4 and 5).

3.3.2 Local Monetary Policy Rate and Bank Credit: Carry Trade Mechanism

In this subsection, we investigate a mechanism for the previously documented results. In particular, we ask whether CC interact with the relation between local policy rates and bank debt, while controlling for the potential simultaneous effects of RR-taxed liabilities. We first present the empirical strategy and next discuss the findings.

EMPIRICAL STRATEGY

We present the most robust version of the model, estimated over the period 2005Q2-2008Q2:

$$
\begin{aligned}
Y_{f,b,yq} = {} & \text{FX-Funds}_{b,yq-1} * \left(\beta_1 + \beta_2 \text{MPspread}^{US}_{yq-1} + \beta_3 \text{Post}_{yq} + \beta_4 \text{MPspread}^{US}_{yq-1} * \text{Post}_{yq} \right) + \\
& \text{SavingD}_{b,yq-1} * \left(\sigma_1 + \sigma_2 \text{MPspread}^{US}_{yq-1} + \sigma_3 \text{Post}_{yq} + \sigma_4 \text{MPspread}^{US}_{yq-1} * \text{Post}_{yq} \right) + \\
& \text{CheckingD}_{b,yq-1} * \left(\phi_1 + \phi_2 \text{MPspread}^{US}_{yq-1} + \phi_3 \text{Post}_{yq} + \phi_4 \text{MPspread}^{US}_{yq-1} * \text{Post}_{yq} \right) + \\
& \text{FX-Funds}_{b,yq-1} * \left(\mu_1 + \mu_2 \text{Macro}_{yq-1} + \mu_3 \text{Post}_{yq} + \mu_4 \text{Macro}_{yq-1} * \text{Post}_{yq} \right) + \\
& \text{BankControls} * \left(\Gamma_1 + \Gamma_2 \text{MPspread}^{US}_{yq-1} + \Gamma_3 \text{Post}_{yq} + \Gamma_4 \text{MPspread}^{US}_{yq-1} * \text{Post}_{yq} \right) + \\
& + \delta_{f,b} + \delta_{f,yq} + \varepsilon_{f,b,yq}
\end{aligned}
$$

The dependent variable, $Y_{f,b,yq}$, is the log total volume of outstanding debt provided by bank b to firm f in year:quarter yq (i.e. $\text{Loan}_{f,b,yq}$). We study how these variables react to variations in the local versus

U.S. policy rate spread, depending on the banks' relative reliance on FX liabilities (affected by the CC). To better highlight the carry trade mechanism, we report in our main table results employing the policy rate spread, but we also show that results are robust if we use the simple lagged local policy rate.

The main coefficients of interest are β_2 and β_4. Under the carry trade hypothesis, β_2 is positive, as banks with higher FX funding lend more when the wedge between the policy rates goes up, while β_4 is negative, as CC break the carry by increasing the costs of FX funding, thereby reducing the gains associated with larger policy rate wedges. Crucially, we horse race our carry trade mechanism against the alternative hypothesis that domestic deposit funding drives the different relation between loan volume and local monetary policy rate before and after 2007Q2. In such a case, RR are important instead of CC for the strengthening of the bank-lending channel of monetary policy, and the inclusion of the interaction of savings and checking deposits with both the policy rate spread and the post dummy nullifies the coefficients β_2 and β_4.[92]

We control as well for the interactions of bank FX funding with the other macro controls (and with the post-dummy), e.g. because variations of the exchange rate, which correlate with the MP-spread, might induce a different reaction in credit supply across differently FX exposed banks. Moreover, we allow for all remaining bank characteristics to influence bank debt differently depending on the lagged level of the policy rate spread, before and after the enforcement of the policy (e.g. higher levels of bank capitalization are associated with credit expansions when the interest rate is relatively higher, Jiménez et al., 2012).

Finally, we saturate the model with firm*bank fixed effects, $\delta_{f,b}$, and, importantly, firm*year:quarter fixed effects, denoted by the parameters $\delta_{f,yq}$. Following Khwaja and Mian (2008) and Jiménez et al. (2012 and 2014), these fixed effects are crucial for the isolation of the bank lending channel of monetary policy, as they allow the comparison of the evolution of credit to the same firm in a given year:quarter in reaction to variations of the policy rate, depending on the different funding structures of the firm's lenders. In other terms, such fixed effects fully control for firms' time-varying demand shocks.

[92] Formally, the simple sensitivity of β_2 and β_4 to the inclusion of the full interaction of domestic deposits with the interest rate spread and the post dummy does not prove itself that RR are key to strengthening the bank-lending channel of monetary policy. This would also require that σ_2 and ϕ_2 are positive, so that savings and checking deposits drive the positive association between policy rate variations and loan size before 2007Q2, and that σ_4 and ϕ_4 have negative values, suggesting that shocks to RR are responsible for the more negative correlation between loan volume and variations in the local policy rate after the enforcement of the measures.

RESULTS

We report results in Table 4. Note that, with respect to the largest sample in Table 2, the inclusion of firm*year:quarter fixed effects reduces sample size, as it excludes all companies with just one lending relationship. However, we validate the baseline results from the previous subsection (i.e. using time variation only) in this smaller sample in column 1.

In column 2, we start testing the carry trade mechanism by fully interacting bank FX funding with the lagged interest rate spread and the post dummy. We also control for the full interaction of banks FX funding with the other macro controls, and allow other bank characteristics to exert a different, unconditional, impact on credit before and after the introduction of the prudential measures in 2007Q2. Two results emerge immediately. First, in line with carry trade lending strategies, β_2 is positive and β_4 is negative. That is, before CC, banks with higher share of FX funds expand credit relatively more when the spread goes up, while after CC they reduce lending in reaction to a positive variation of the spread. Importantly, $|\beta_4| > |\beta_2|$, suggesting that CC revert the dynamics, instead of just attenuating it, thereby contributing to restoring the transmission of local interest rate policy to bank credit.

In column 3, we control for time-varying macro-economic shocks by including year:quarter fixed effects and coefficients are virtually unaffected. In column 4, we take a more serious step in the direction of isolating credit supply channels by augmenting the model with industry (four-digit SIC)*year:quarter shocks, but the resulting variation in coefficients is again minimal. In column 5, we finally introduce firm*year:quarter fixed effects, therefore fully controlling for time-varying firm idiosyncratic demand shocks. If anything, the magnitude of the coefficients β_2 and β_4 increases.

Finally, in column 6, we report the most robust version of the model where we additionally interact all the banks characteristics with the policy rate spread and with the post dummy. Validating the carry trade mechanism, β_2 and β_4 are further strengthened. Importantly, the impact of carry trade lending strategies on bank domestic credit is both statistically and economically significant. In reaction to a 1 p.p. jump in the policy rate spread, before CC, banks with a 1 s.d. (i.e. 2.6 p.p., see Table 1) higher share of FX funds expand credit in relative terms by 3.8 p.p.. After CC, however, the same combination of spread-increase and larger FX funding is associated with a relative reduction in credit supply by 3.5 p.p.. The application of CC therefore sharply reduces carry trade incentives and contributes to restoring a negative relation between local policy rate variations and credit among highly FX indebted banks.

Differently, as shown in the Table A3 in the Appendix, whereby we display the horse race between the FX and the RR taxed liabilities, the interaction of the latter domestic liabilities with the policy rate spread (and the post dummy) is not significant, suggesting that the change in RR-policy did not contribute to strengthening the bank-lending channel of monetary policy rates.

ROBUSTNESS AND ADDITIONAL FINDINGS

In Table 5, we estimate alternative specifications of the model. In column 1, we further control for loan loss provisions, and their interaction with both the policy rate spread and the post dummy. In column 2, we rerun the baseline model (in column 6 of Table 4) with observations weighted by log loan size, so to allow our coefficients to be driven from more meaningful credit relationships. In column 3, we complement the WLS estimation with the full interaction of the loan loss provisions with the policy rate spread and the post dummy. In columns 4 to 6, we estimate both by OLS and WLS the baseline model (and its augmented version with loan loss provisions), removing firm*year:quarter fixed effects and substituting them with industry*year:quarter fixed effects. This allows us to retrieve information on bank credit for companies borrowing from one bank only. Importantly, coefficients are virtually unaffected from all such modifications of our baseline model, so that both the qualitative and quantitative interpretation of our channel provided in the previous subsection go through.[93]

We estimate the baseline model under alternative clustering strategies at the level of firm and bank in column 7, firm and bank and year:quarter in column 8, firm and bank*industry and year:quarter in column 9, and find that our coefficients of interest are nonetheless significant at conventional levels.

We next examine the evolution of the carry trade lending mechanism over our period of analysis. In our regressions, we check the relative difference of the conditional response of FX indebted banks to the policy rate spread before and after the introduction of CC in 2007Q2. However, one might worry that the contraction in the strength of carry trade lending strategies that we attribute to the CC period might reflect a declining trend that took place before the introduction of CC. To address such concerns, we estimate the following model:

$$Loan_{f,b,yq}=(\beta_1+\beta_2 Pre_{yq}+\beta_3 Post_{yq})*MPspread^{US}_{yq-1}*FX\text{-}Funds_{b,yq-1}+Controls_{b,yq-1}+\delta_{f,b}+\delta_{f,yq}+\varepsilon_{f,b,yq}$$

where: Pre_{yq} is a dummy with value 1 from 2006Q2 onward and 0 otherwise, whereas $Post_{yq}$ is our usual post-policy dummy with value 1 from 2007Q2 onwards.[94] Under this specification: β_1 represents the intensity of the carry trade lending strategy during the period 2005Q2 to 2006Q1; β_2 estimates the

[93] We further validate that our findings are robust to controlling for the full interaction of the policy rate spread and the post-dummy with an indicator for whether a bank is foreign-owned or not (the related regression table is available upon request). Indeed, our main coefficients of interest related to the carry-trade channel remain both qualitatively and quantitatively unchanged.

[94] $Controls_{b,yq-1}$ reflects the most robust version of the model, in which other balance sheet items are also fully interacted with $MPspread^{US}_{yq-1}$ (and both the Pre_{yq} and the $Post_{yq}$ dummies) and $FX\text{-}Funds_{b,yq-1}$ is further interacted with macro controls (and both the Pre_{yq} and the $Post_{yq}$ dummies). $\delta_{f,b}$ and $\delta_{f,yq}$ denote firm*bank and firm*year:quarter fixed effects.

intensity over 2006Q2-2007Q1 compared with the previous period; and finally, β_3 measures the change in the strength of carry trade lending over the CC-period 2007Q2-2008Q2 (relative to the period 2006Q2-2007Q1). We depict these three coefficients in Figure 3. The coefficient β_1 is positive but not statistically significant, whereas the larger and statistically significant coefficient β_2 suggests that the strongest period for carry trade has been 2006Q2-2007Q1.

Finally, β_3 is strongly negative and statistically significant, ultimately suggesting that the contraction in carry trade lending before the policy cannot be attributed to preexisting trends. Moreover, the strengthening of the carry trade lending strategy over the period 2006Q2 to 2007Q1 is consistent with both aggregate and bank-level figures on FX inflows. Note in Figure 2, Panel B, it is around 2006Q2 that Colombia has large (non-FDI) capital inflows associated with a higher policy rate which widens the spread with the U.S. Effective Federal Funds Rate. In Figure 4, we show the total quarterly FX debt intakes by Colombian banks (through long-term loans and bonds).[95] Also in this more granular chart the capital boom ramps up at the end of 2006 and beginning of 2007 and is eventually halted by CC, therefore tightly mirroring the dynamics portrayed by our estimates.

An especially important and interesting set of controls in our analysis is provided by the interaction of bank FX funding with proxies of conditions in the Colombian external sector, which might influence the ability of local banks to access FX liabilities as well as the value of such liabilities throughout time. In Table 6, we show our findings under alternative proxies. First, in column 1, we display the baseline model in column 6 of Table 4, in which we use the log exchange rate (expressed as Colombian Pesos per 1 USD, so that an increase denotes a depreciation of the local currency). Interestingly, before CC, a 1 s.d. appreciation of the exchange rate triggers a relatively larger increase in lending by 1 p.p. among banks with a 1 s.d. larger FX funding. However, under CC, this effect is not significant. In column 2, we replace the log exchange rate with the VIX, commonly interpreted as an indicator of global risk aversion (liquidity conditions) that significantly responds to U.S. monetary policy shocks (Miranda-Agrippino and Rey, 2020) and drives capital flows worldwide (Rey, 2015).[96] Our main coefficients of interest are robust to this replacement. Moreover, the interaction of the VIX with banks FX funding suggests that a 1 s.d. loosening in global risk aversion (decline in VIX) is associated with a relative jump in lending by 2.4 p.p. for banks with a 1 s.d. larger FX funding pre-CC. In line with results for the exchange rate, though, such influence is nullified by CC. Finally, in column 3, we use oil price as an indicator of external sector conditions for Colombia. Despite oil representing the bulk of Colombian

[95] Importantly, they exclude FX-liabilities issued by Colombian banks through foreign subsidiaries and therefore significantly underestimates the extent of FX-borrowing.

[96] Reflecting a significant interdependence between the VIX indicator and external sector conditions in Colombia, the joint inclusion of the VIX and the exchange rate in a regression model generates multicollinearity issues. For this reason, we include the two variables in alternative models rather than together. Similar considerations apply to oil price, which is the main driver of Colombian exchange rates, given the prominent role of oil exports.

exports, its price is largely determined by exogenous factors and comoves substantially with the exchange rate (over our period, by a factor of 80 percent). Once again, the carry trade coefficients are not affected and a 1 s.d. increase in oil prices drives a relative expansion of lending by 5 p.p. for banks with a 1 s.d. larger FX funding, an influence halted by CC. Overall, these results are consistent with a mechanism such that when global conditions are loose, the value of bank FX liabilities increases and so does their credit supply (see, e.g., Bruno and Shin, 2015b). This channel is eventually broken by CC.

Finally, we check in Table A4 in the Appendix that the carry trade channel is robust to substituting the policy rate spread with the simple policy rate. The coefficients from the most robust version of the model in column 6 imply that, before CC, banks with a 1.s.d. larger FX funds respond to a 1 p.p. jump in the policy rate by increasing lending by 1.8 p.p.. Under CC, the same combination of policy rate increase and larger FX funding is associated with a relative lending cut by 1.5 p.p.

DISSECTING THE CARRY TRADE MECHANISM: PESO *VS* FX-LENDING AND CIP DEVIATIONS

In Table 7, we repeat the exercise over a smaller sample of observations for which loan volume is broken down by currency, i.e. domestic (peso) and FX-lending.[97] We consider peso and FX-loans separately in column 1 and 2, respectively. The coefficients clearly show that peso lending drives the results just described, which further corroborates the carry trade hypothesis.

While a carry trade strategy is in principle profitable also with FX loans (under the reasonable assumption that these are more expensive in Colombia than in global interbank and wholesale funding markets), the strategy will nonetheless grant higher returns through peso-lending, given the positive policy rate differential. Also, banks would bear the CC tax only if FX funds were reinvested in peso-denominated assets, so that credit supply variations induced by CC must show up among peso loans. If banks borrow in FX to finance FX loans, the CC is borne by the ultimate borrower, i.e. a local company. As a result, the reduction in FX-loans associated with CC should show up as a demand shock in a loan level analysis, but we fully control for it by adding firm*year:quarter fixed effects (Khwaja and Mian, 2008). Finally, in column 3, the share of peso loans out of total bank debt also evolves according to the carry trade lending, although pre-policy carry is just marginally significant, and column 4 replicates the baseline analysis on total bank debt in this smaller sample by summing up peso and FX loans.

Moreover, we investigate the relation between the documented carry trade strategies by local banks and the deviations from CIP, computed over the 3-month yield spread between Colombia and U.S. sovereign bonds. Importantly, banks tend to fully hedge their FX borrowing. Hence, validating that our results are

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[97] We report the summary statistics for this smaller sample in Table A1 of the Appendix. Note that this sample consists of large companies with supervised balance sheets and accounts across time for 55 percent to 60 percent of the total loan volume for multibank nonfinancial companies in the regression sample in Table 4.

robust to substituting the policy rate spread with a proxy for CIP deviations is relevant, because such deviations grant returns from the carry even under fully hedged currency risk.

In Table 8 we report results from this exercise. We start by substituting the policy rate spread with the spread between the Colombian and U.S. sovereign 3-month yields. As already detailed in the data section, the two variables are very tightly linked, and in practical terms the sovereign spread corresponds to the policy rate differential plus a half p.p. premium. We repeat our regressions with the sovereign spread since the CIP-deviations retrieved from Du and Schreger (2016) are based on it. Indeed, results in column 1 of Table 8 confirm that carry trade lending strategies are operative also based on such 3-month sovereign spread. Quantitatively speaking, before the introduction of CC, a 1 p.p. increase in the sovereign spread triggers a relative jump in credit supply of 2 p.p. by banks with a 1 s.d. higher FX funding. After the introduction of CC, however, the same combination of sovereign spread hike and higher banks FX funding leads to a 2.76 p.p. cut in credit supply.

In column 2, we introduce the deviations from the CIP. Again, the coefficients are consistent with carry trade lending. That is, before (after) CC, higher CIP-deviations bring relatively larger (smaller) volumes of credit supply for more FX-exposed banks. A reasonable concern is that these results are confounded by the fact that CIP-deviations tend to be higher at times of relatively high sovereign spread, so that they do not necessarily reflect carries prompted by returns under full hedging of currency risk. For this reason, we perform an additional regression in column 3 in which we augment the model with the component of the 3-month sovereign spread which is not accounted for by CIP-deviations, namely the 3-month forward premium, $FP_{yq-1.}$ We perform a full horse-race in which also this factor is fully interacted with not only bank FX funding, but also with the other bank controls. The resulting estimates suggest that the carry trade lending strategy is driven by both forward premia and CIP deviations. Nonetheless, CIP deviations exert a relatively greater influence on the dynamics of credit supply for FX indebted banks, suggesting that banks are relatively more inclined to pursue fully hedged carries (against FX risk). Absent CC, among banks with a 1 s.d. higher FX funding, a 1 p.p. increase in CIP deviations (forward premia) triggers an increase in credit by 3.5 p.p. (1.7 p.p.). With CC in place, the same combination of higher CIP deviations (forward premia) and bank FX funding leads to a 7.7 p.p (1.3 p.p.) decline in credit. Finally, we replicate the analysis in column 4 with peso loans, in column 5 with FX loans and in column 6 with the share of peso loans as dependent variables. Consistent with our previous findings, higher CIP deviations (and forward premia as well, but less intensively) drive carry trade predominantly over peso loans for banks with relatively larger FX funding.

CARRY TRADE MECHANISM AND RISK-TAKING: HETEROGENOUS EFFECTS ACROSS FIRMS

We investigate whether carry trade lending heightens bank risk-taking, and the eventual influence of CC on it, by looking for heterogenous effects across companies, depending on their riskiness and opaqueness. In detail, we proxy for credit risk by sorting companies based on quartiles of the distribution of the average interest payments in the pre-policy period, i.e. Firm $Risk_{f,pre.}$ An identical classification ranks firms by liquidity risk based on the distribution of the average pre-policy reliance on short-term debt, i.e. Short-Term $Debt_{f,pre}$. Additionally, we split companies depending on whether they defaulted on at least one loan during the period 2005Q1-2007Q1, which further proxies for default risk. Finally, we divide companies by transparency and opaqueness based on whether their balance sheet is publicly supervised or not.

Table 9 reports estimates from the regressions on loan volume from the most robust version of the model. To start with, companies in the first quartile of credit risk (Firm $Risk_{f,pre}$) are not impacted by the carry trade, neither before nor after CC. On the contrary, firms with greater credit risk experience larger fluctuations in bank debt associated with the carry, and especially so for companies with above-median credit risk. For instance, in reaction to a 1 p.p. jump in the policy rate spread, before CC, banks with a 1 s.d. higher share of FX funds expanded credit to firms in the fourth quartile of credit risk by 7.3 p.p. After CC, these firms also suffer sharper cuts, by 7.7 p.p.

Similar dynamics apply to firms with different levels of liquidity risk. Indeed, carry trade lending does not affect bank debt of firms with the lowest liquidity risk, but significantly impacts loans to firms with higher reliance on short-term debt. Moreover, before CC, when the spread goes up by 1 p.p., companies that ex-ante default (do not default) on one or more loans enjoy a relative credit expansion of 5.1 p.p. (3 p.p.) by banks with a 1 s.d. higher share of FX funds; after CC, the same combination of jumps in the spread and in lenders' FX funds brings a relative credit reduction of 4.7 p.p. (3.2 p.p.), suggesting that both the credit expansion due to the carry and the CC-induced cut are stronger among riskier companies. Finally, opaque firms do not benefit more than transparent ones from carry lending before CC, but after their enforcement they undergo a much larger reduction in credit, by 5.5 p.p. in response to the usual 1 p.p. increase in the spread and 1 s.d. jump in banks' FX exposure.

Overall, these findings are consistent with increased bank risk-taking due to carry trade lending. Also, they indicate that CC contribute to mitigating these risks, as the post-CC reduction in lending by highly FX indebted banks (following an interest rate spread increase) is concentrated among risky and opaque borrowers.

3.4 The impact of Reserve-Requirements on Bank Credit

In this section, we evaluate the impact of the shocks to RR on domestic deposits on bank credit. First, we present the empirical strategy. Second, we discuss the baseline findings. Third, we provide some robustness checks. Fourth, we check whether the impact of RR is heterogeneously distributed across firms. Last, we ask whether bank domestic deposits and foreign funding are substitutes or complements,[98] which reveals whether RR and CC affect credit supply through distinct channels.

3.4.1 Empirical Strategy

We run a difference-in-differences exercise in symmetric five-quarter windows around the modification of the RR-policy in 2007Q2, i.e. over the period 2006Q1 to 2008Q2. We employ the following model:

$$\text{Loan}_{f,b,yq} = \beta_1 \text{Post}_{yq}\text{*RR-Depo}_{b,2007Q1} + \gamma \text{Post}_{yq}\text{*BankControls}_{b,2007Q1} + \delta_{f,b} + \delta_{f,yq} + \varepsilon_{f,b,yq}$$

The dependent variable is loan volume. The main coefficient of interest is β_1, describing the impact of ex-ante heterogeneity in RR-taxed deposits, i.e. the sum of checking and saving deposits, on the ex-post volume of credit. Note that heterogeneity across bank reliance on deposits taxed by the RR-shock is taken as of 2007Q1, and an identical convention is applied to bank controls. The model is augmented with firm*bank dummies, whereas firms' credit demand is controlled through firm*year:quarter fixed effects (Khwaja and Mian, 2008). $\varepsilon_{f,b,yq}$ is an error term, double-clustered at the firm and bank*industry level.

The consistency of our estimates crucially depends on the parallel trend assumption: absent the modification of RR policy in 2007Q2, banks with different reliance on checking and savings deposits would experience parallel ex-ante and ex-post credit dynamics. We test the validity of such assumption in our setting using the alternative model:

$$\text{Loan}_{f,b,yq}=\beta_{1,yq}\text{*(1[year:quarter=yq])*RR-Depo}_{b,2007Q1} + \gamma \text{Post}_{yq}\text{*Bank Controls}_{b,2007Q1} + \delta_{f,b} + \delta_{f,yq} + \varepsilon_{f,b,yq}$$

That is, we allow the relation between RR-Depo$_{b,2007Q1}$ and loan volume to vary over the different year:quarters in our sample, as 1[year:quarter=yq] is a dummy variable with value 1 in year:quarter yq and 0 otherwise. We fix 2006Q1 as the baseline period. A heuristic validation of the parallel trend assumption requires that the RR-treatment effect is zero before 2007Q2, and significant thereafter.

3.4.2 Baseline Results

We report baseline results in Table 10. In column 1, we apply a minimal set of controls, including firm fixed effects and bank controls, interacted with the post dummy. The treatment effect is negative and significant at the 1 percent level. We next saturate the model, first by including firm*bank fixed effects, which imply an increase in the R-squared by roughly 30 p.p.. The treatment effect remains negative and significant (column 2). We then control for time-varying shocks, either common across all firms (column

[98] Note that the sum of domestic (saving and checking) deposits and FX-funding constitutes, on average, roughly 54 percent of banks' total assets. Hence, whether banks with a higher share of domestic deposits are more or less indebted in FX is ultimately an empirical question.

3) or industry-specific (column 4), by applying year:quarter and industry*year:quarter fixed effects, respectively. Coefficients are virtually unaffected.

In column 5, we fully shut down firm demand shocks with firm*year:quarter fixed effects. The treatment effect remains significant at the 1 percent level, suggesting also a strong economic impact of RR-shock. A 1 s.d. (7.8 p.p.) increase in the share of total assets financed with either savings or checking deposits implies a 5.4 p.p. reduction in bank credit. In column 6, we test separately for the effect of checking and savings deposit exposures. The coefficients suggest a stronger effect of exposure to checking deposits, as a 1 s.d. (4.1 p.p.) jump implies a 7.6 p.p. reduction in loan volume. The effect of exposure to saving deposits is smaller, but nonetheless economically meaningful, corresponding to about 2.9 p.p. in reaction to a 1 s.d. (7 p.p.) increase.

3.4.3 Robustness

First, in Panel A of Table 11, we estimate alternative specifications of the model. In column 1, we further control for loan loss provisions, both alone and interacted with the post dummy. In column 2, we rerun the baseline model (from column 6 of Table 10) with observations weighted by log loan size to allow our coefficients to be driven from relatively larger credit relationships. In column 3, we complement the WLS estimation with loan loss provisions and their interaction with post dummy. In columns 4 to 6, we estimate, both by OLS and WLS, the baseline model (and its augmented version with loan loss provisions), removing firm*year:quarter fixed effects and substituting industry*year:quarter fixed effects. This allows us to include in the regression sample those companies that borrow from only one bank. Importantly, coefficients are virtually unaffected by all such modifications of our baseline model, so that both the qualitative and quantitative interpretation of our channel provided in the previous subsection go through.

In columns 7, 8, and 9, we estimate the baseline model under alternative clustering strategies. In column 7, we estimate at the level of firm and bank. In column 8, firm and bank and year:quarter. And in column 9, firm and bank*industry and year:quarter. Our coefficients of interest remain significant at least at the 12 percent level in the case of firm and bank clustering.

Second, we run cross-sectional regressions in which the dependent variable is the loan growth rate between 2007Q1 (the year:quarter before the shock to RR) and j quarters ahead, j={1,2,3,4,5}, to validate that the negative treatment effect persists across periods shorter than the five-quarter period we consider in the baseline finding. Results in Panel B of Table 11 suggest that this is the case.

Third, we further inspect the validity of the parallel trend assumption. Figure 5 depicts the time-varying coefficient of the treatment effect (relative to a baseline, fixed at zero, for 2006Q1). Indeed, before 2007Q2, overall exposure to savings and checking deposits does not affect bank credit. After the RR shock, however, the coefficient becomes markedly negative and statistically different from zero, which provides suggestive evidence in favor of the parallel trend assumption being verified.

Fourth, we run a placebo test. That is, we consider a pre-policy sample from 2005Q1 to 2006Q4, and fix exposures and bank controls as of 2005Q4. This is a "fake" exposure, which should not be associated with a contraction in credit, which is confirmed in Panel C of Table 11.[99]

3.4.4 Heterogenous Effects across Firms

As with carry trade regressions, we sort companies according to proxies of credit risk, liquidity risk, default risk, and opaqueness, and repeat the baseline exercise across such different groups of firms. Table 12 displays the results.

The reduction in credit is not significant among firms with the lowest credit risk (those in the lowest quartile of the ex-ante distribution of average interest payments over loans). On the other hand, it is significant across riskier companies, and the reduction in credit among them increases as their riskiness does. In particular, firms in the upper quartile of credit risk experience a 15.4 p.p credit on loans from banks more RR-exposed by a 1 s.d. increase. Similarly, only companies with above-median liquidity risk suffer credit reduction from more RR-exposed financial institutions. Furthermore, there is not a statistically significant difference between companies with and without ex-ante loan defaults, but stark differences emerge between transparent companies and opaque companies. The former do not suffer any credit reductions due to RR shocks, whereas the latter suffer a 7.8 p.p. credit cut by lenders more exposed to RR by a 1 s.d. increase.

3.4.5 Banks' Domestic and Foreign Funding: Complements or Substitutes?

We have so far shown that: i) the bank-lending channel of monetary policy rates is strengthened by CC, which affects bank foreign liquidity; ii) the shocks to RR exert a large direct negative effect on bank credit by raising the cost of core domestic liquidity. Both policies therefore contribute to taming credit booms. It remains to be understood whether the foreign and domestic liquidity are complements or substitutes in bank funding structure, i.e. whether banks that use more FX funds also employ larger core deposits to finance their assets, or not.

The scatterplot in Figure 1, which reports bank (time-varying, quarterly) reliance on savings and checking deposits on the x-axis and bank FX funds on the y-axis, indicates that banks that use more FX liquidity rely less on domestic core deposits. In other terms, banks more exposed to RR are *less* exposed to CC, and over the period of analysis the two variables are correlated negatively (by a factor of 37

[99] The summary statistics for the placebo test are in Table A2 of the Appendix.

percent, significant at 1 percent level). A formal way to discern whether RR and the CC operate independently from each other is to directly horse race them in a regression model. In Table 13, we show results from such an exercise (run over the longer period from 2005Q2 to 2008Q2), in which we contemporarily employ the full interaction of the policy rate spread with banks FX funds and the post dummy, as well as the full interaction of the RR-taxed checking and savings deposits with the post dummy.[100] In those regressions, both the decline in carry trade lending due to CC and in credit provided by banks more reliant on RR-taxed liabilities are significant, suggesting the two macroprudential policies operate independently from each other.

Put differently, CC and RR - i.e. macroprudential measures targeting foreign and domestic bank debt, respectively - affect bank credit supply through different channels, as banks more affected by CC are less impacted by RR, and vice versa. Both measures are therefore needed to slow down a boom driven by both foreign and domestic liquidity.

3.5 Conclusions

We analyze the effects of capital controls and macroprudential policy on credit supply exploiting: (i) the simultaneous introduction of capital controls and an increase of reserve requirements on domestic bank deposits in Colombia during a strong credit boom; and (ii) administrative credit registry and supervisory bank balance sheet data. In brief, we find the following robust results: first, banks use cheaper FX funding from abroad to arbitrage higher local monetary policy rates (which raises the policy rate spread against the U.S.), by carry trading cheap FX funds with expensive local lending, especially to ex-ante riskier, more opaque local firms. Capital controls, by taxing FX debt, reduce the interest rate differential and break the carry trade, enhancing the bank-lending channel of local monetary (interest rate) policy and reducing bank risk-taking. Second, the increase in reserve requirements on domestic deposits directly reduces credit supply during the boom, and more so for riskier firms, rather than (indirectly) enhancing the effects of monetary policy rates on credit supply.

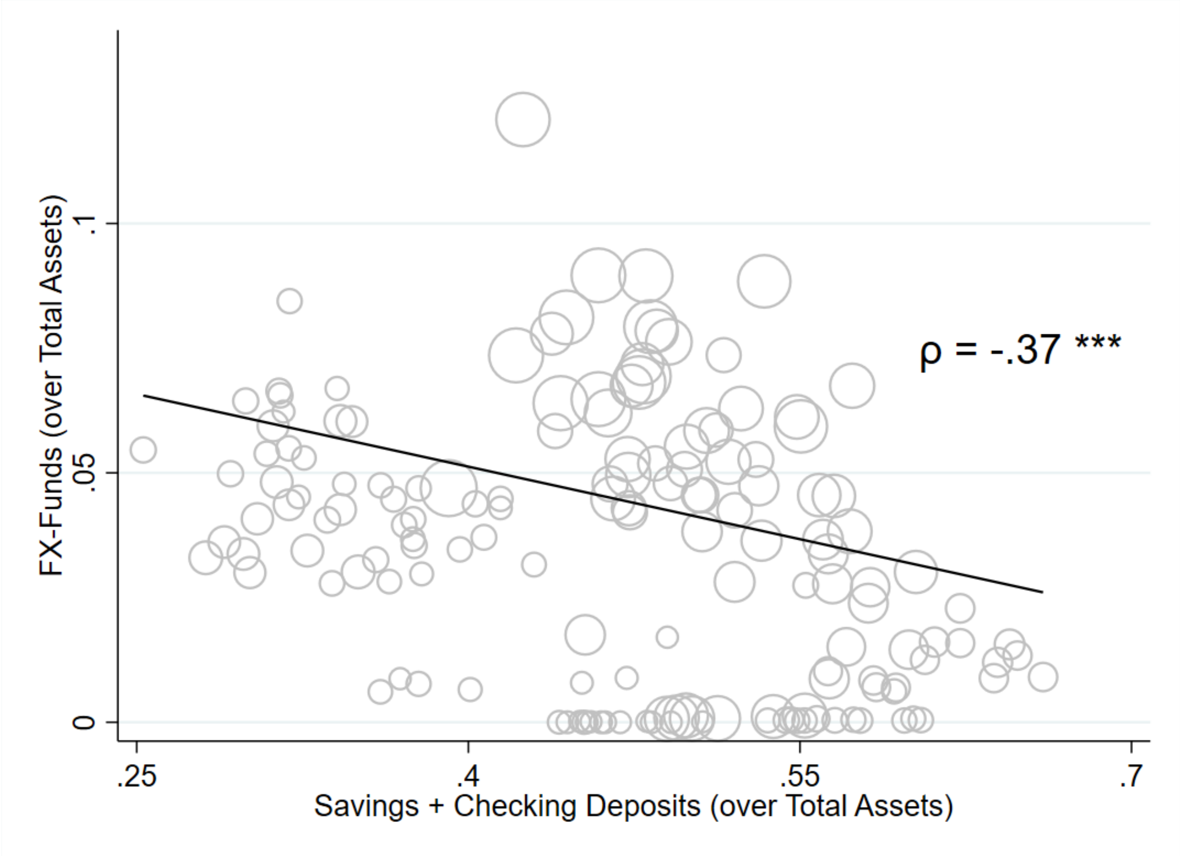
Our main contribution to the literature is to show that both capital controls and (domestic) macroprudential policy tame credit supply booms, including credit supply to ex-ante riskier firms, by targeting different sources of bank debt. Capital controls target foreign bank debt, thereby improving the effectiveness of the bank lending channel of (local) monetary policy -by halting carry trade lending strategies by local banks- and domestic macroprudential policy targets local bank debt, directly attenuating credit supply booms. As credit booms stem from *both* foreign *and* local liquidity, and we find that reliance on domestic deposits *versus* foreign (FX) debt are very negatively correlated across

[100] Note that columns 1 through 4 in Table 13 correspond to columns 2 through 5 in Table 4 (check the carry trade coefficients). In Table 13, however, we explicitly show the effect of RR-taxed liabilities, before and after 2007q2.

banks (so that financial intermediaries more affected by capital controls are less impacted by reserve requirements, and the other way around), our results suggest that a Tinbergen rule with two (macroprudential) instruments is necessary to tackle the two (intermediate) objectives (sources of liquidity).

Figures

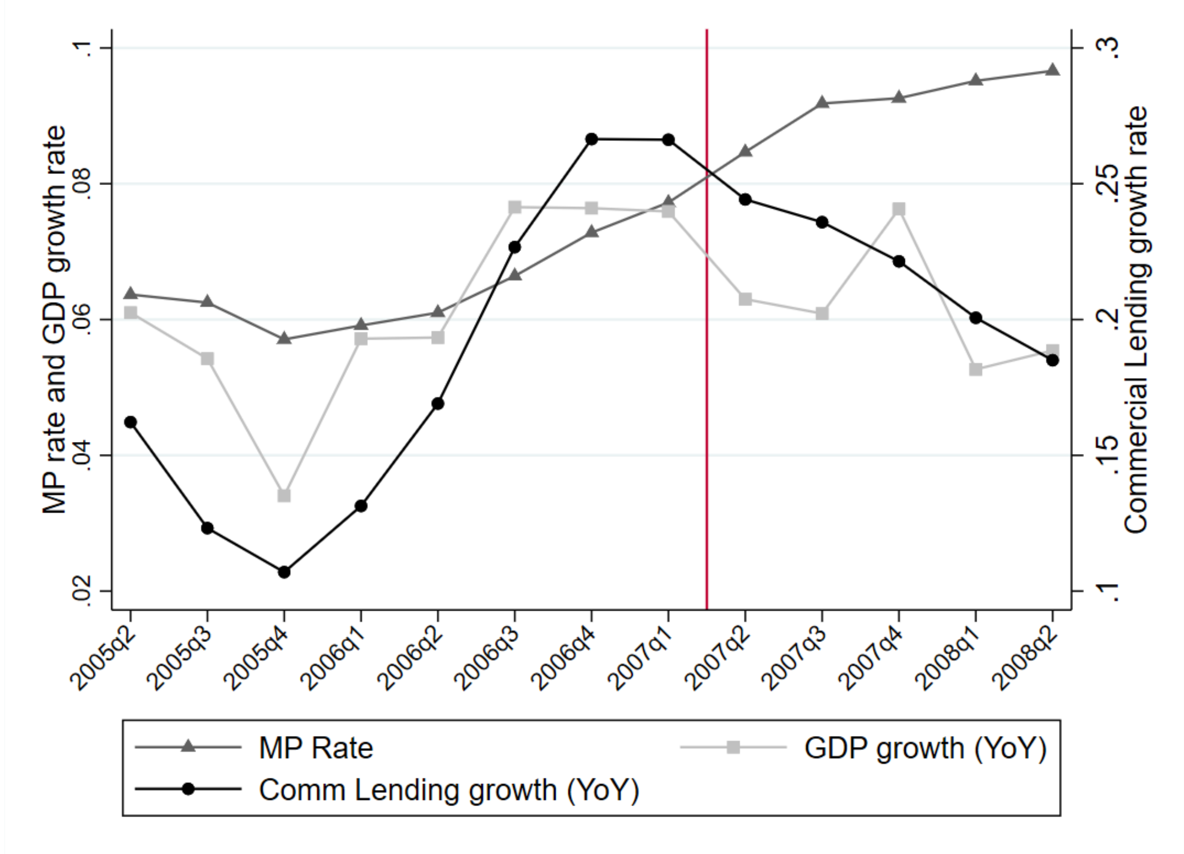
FIGURE 1: Banks FX-Funds *versus* Savings and Checking Deposits



This chart shows the negative correlation between bank FX-funds (y-axis) and bank Savings and Checking Deposits (x-axis) – affected by Capital Controls and Reserve Requirements, respectively – over the period from 2005Q1 to 2008:Q2. Each marker represents a bank-year:quarter pair and is weighted by the relative size (i.e. total assets) of a bank balance sheet with respect to the overall size of the banking sector in a given year:quarter. The coefficient ρ describes the pairwise correlation among the variables, which is equal to -.37 and statistically significant at 1% level.

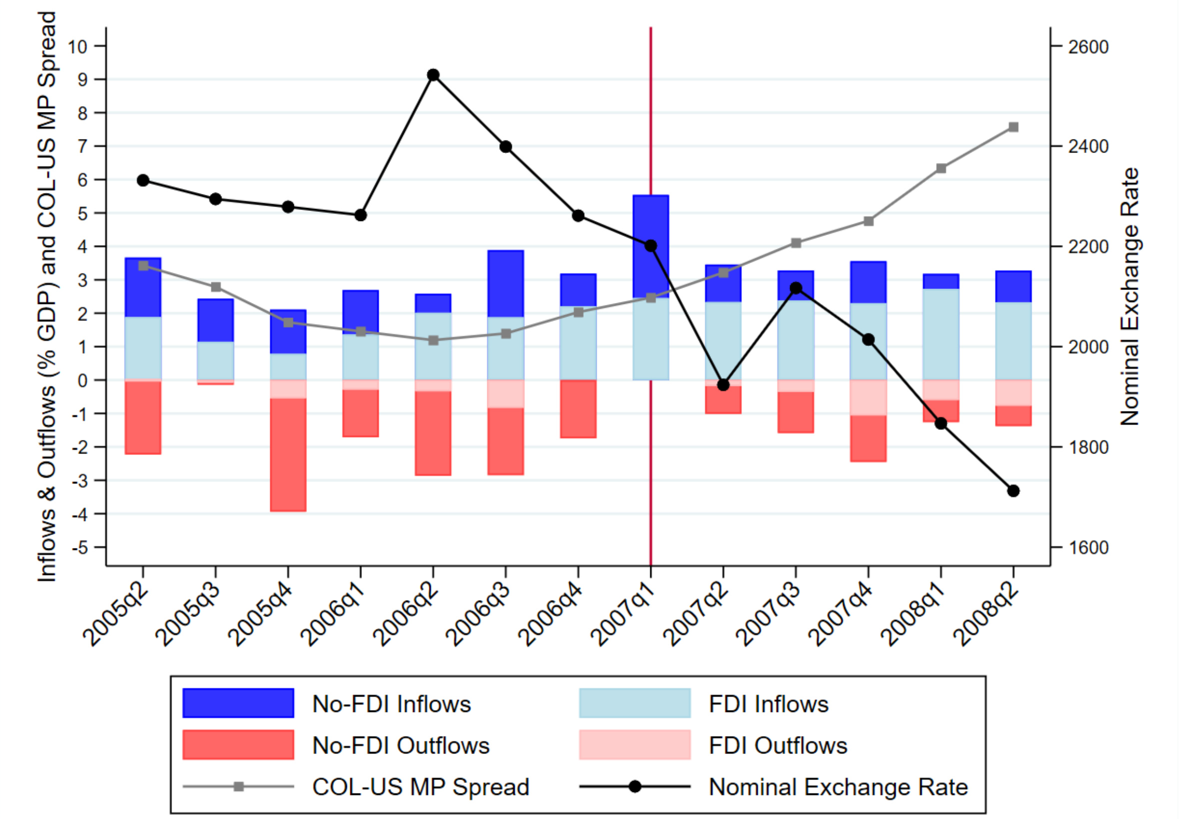
FIGURE 2: Monetary Policy and Credit Growth

Panel A: Credit Growth, Monetary Policy Rate and Economic Growth



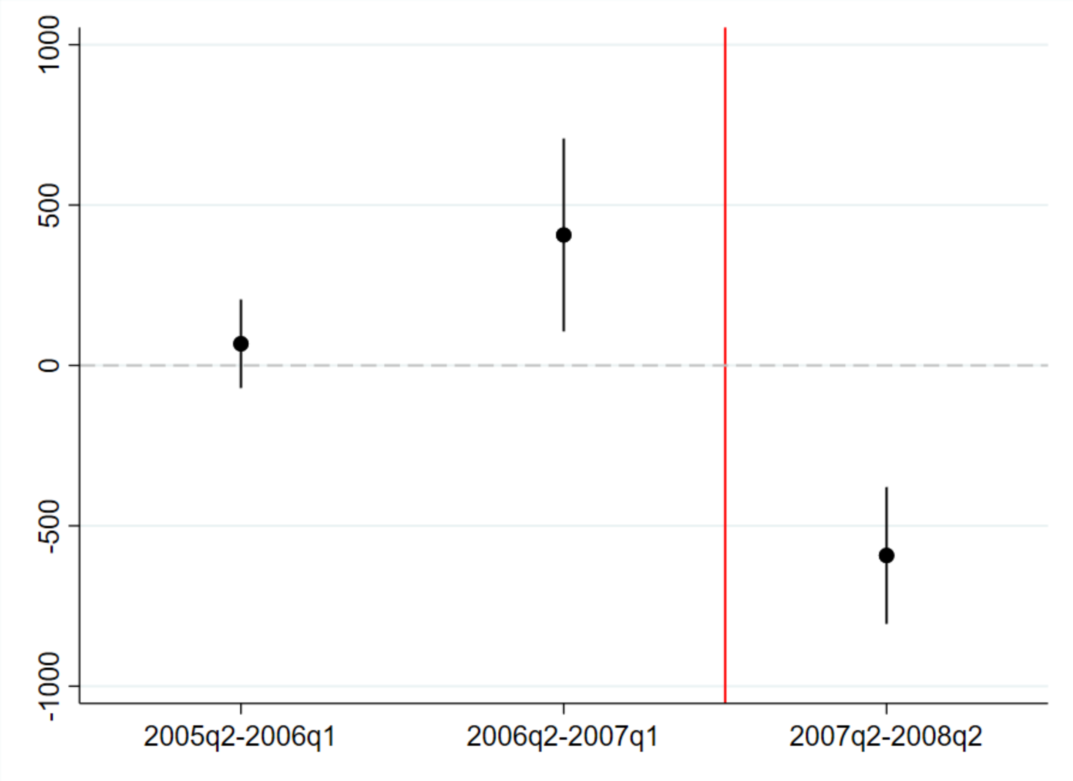
In this figure, the dark gray line – connected by triangles - represents the monetary policy rate (left y-axis), i.e. the prevailing interbank overnight rate. The light gray line – connected by squares - draws the evolution of the yearly growth rate of GDP (left y-axis). The black line – connected by circles - refers to the yearly growth rate of commercial credit (right y-axis). All data are gathered from the Central Bank of Colombia.

Panel B: Exchange Rate, Colombia-US Monetary Policy Rate Spread and Financial Inflows and Outflows



In this figure, the light blue bars represent gross FDI inflows. The dark blue bars denote gross no-FDI inflows, i.e. the sum of gross portfolio inflows and other gross debt inflows. The light red bars represent gross FDI outflows (reported on a negative scale), whereas the dark red bars denote gross no-FDI outflows (also reported on a negative scale). All inflows and outflows measures are expressed as a percentage of GDP on the left y-axis. The gray line - connected by squares - draws the evolution of the spread between the Colombian monetary policy rate, i.e. the prevailing interbank overnight rate, and the Effective FED Funds Rate, expressed in percentage points (left y-axis). The black line – connected by circles - depicts the Colombian Peso/US Dollar nominal exchange rate – i.e. Pesos per 1 US Dollar, so that an increase (decrease) corresponds to a depreciation (appreciation) of the Peso against the US dollar -, measured on the right y-axis. All data are gathered from the Central Bank of Colombia apart from the Effective FED Funds Rate, which are retrieved from FRED.

FIGURE 3: Monetary Policy Rate and Credit: Carry Trade Mechanism over Time

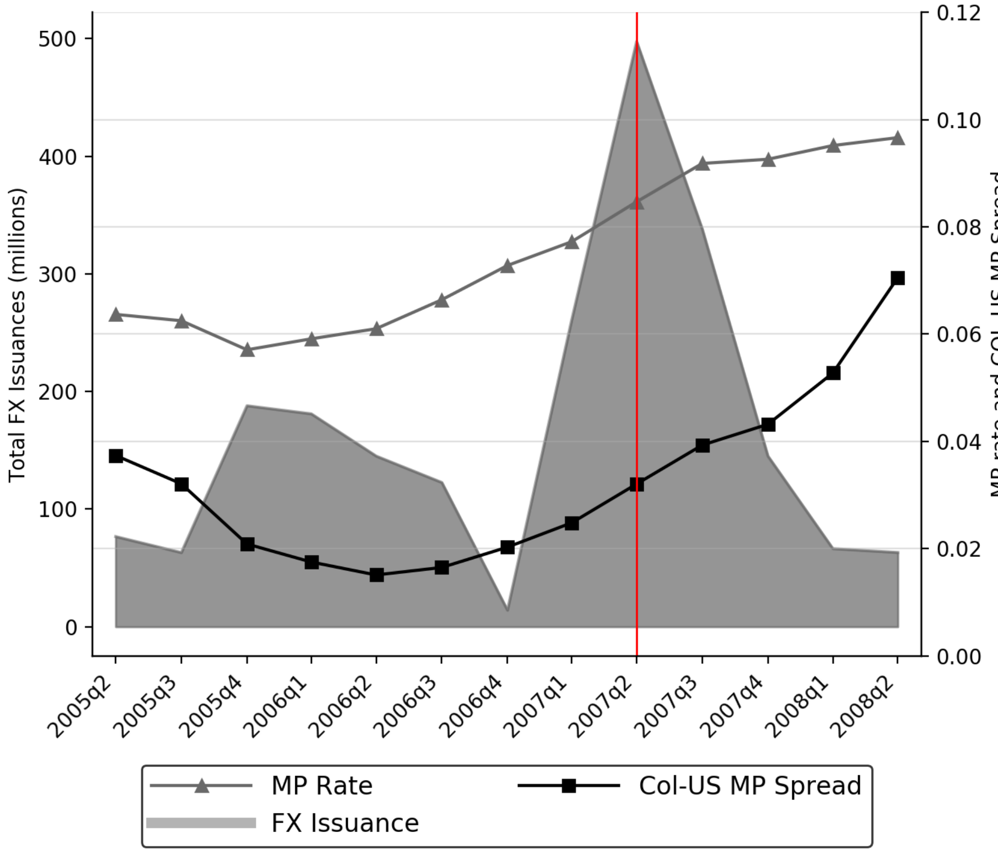


This figure reproduces the time-varying coefficient for the interaction between $MPspread^{US}_{yq-1}$ and $FX\text{-}Funds_{b,yq-1}$ from the following regression:

$$Loan_{f,b,yq}=(\beta_1+\beta_2 Pre_{yq}+\beta_3 Post_{yq})*MPspread^{US}_{yq-1}*FX\text{-}Funds_{b,yq-1}+Controls_{b,yq-1}+\delta_{f,b}+\delta_{f,yq}+\varepsilon_{f,b,yq}$$

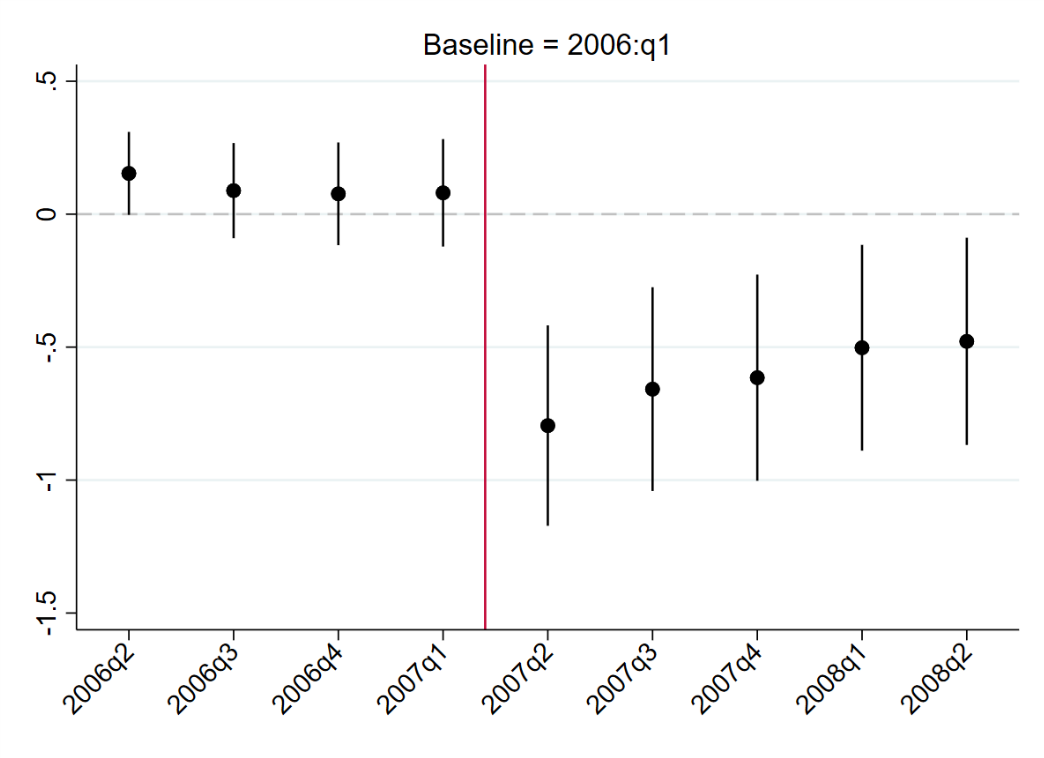
$Loan_{f,b,yq}$ is is the log of total debt provided by bank b to firm f in year:quarter yq. Pre_{yq} is a dummy with value 1 from 2006Q2 onward and with value 0 otherwise. $Post_{yq}$ is a dummy with value 1 from 2007q2 onward and with value 0 otherwise. $Controls_{b,2007Q1}$ is a vector of bank controls (as of 2007Q1) including: ROA, log Total Assets, Common Equity (rescaled by Total Assets), NPL (rescaled by Total Loans) - all being fully interacted with $MPspread^{US}_{yq-1}$ – as well as the full interaction of FX-Funds with the lagged GDP growth rate, CPI index and log exchange rate. $\delta_{f,b}$ is a vector of Firm*Bank fixed effects; $\delta_{f,yq}$ is a vector of Firm*Year:Quarter fixed effects and $\varepsilon_{f,b,yq}$ is an error term. The markers denote the point-estimate of the time-varying coefficients and the lines around them are 95% confidence interval. Standard errors are double-clustered at the firm and bank*industry level.

FIGURE 4: FX Liabilities, Monetary Policy Rate and Colombia-U.S. Monetary Policy Rate Spread



In this figure, the dark gray area represents the total amount of FX liabilities issued by Colombian banks (left y-axis), measured as the two-quarter moving average of total issuances of bonds and long-term loans (excluding issuances by Colombian banks through foreign subsidiaries). The dark gray line shows the monetary policy rate (right y-axis), i.e. the prevailing interbank overnight rate. The black line represents the evolution of the spread between the Colombian monetary policy rate and the Effective FED Funds Rate (right y-axis). All data are gathered from the Central Bank of Colombia apart from the Effective FED Funds Rate, which is retrieved from FRED.

FIGURE 5: Reserve Requirements Shock – Time-Varying Coefficient



This figure reproduces the time-varying coefficient for RR-Depo$_{b,2007Q1}$ (given by the sum of checking and savings deposits as of 2007Q1) from the following regression:

$$\text{Loan}_{f,b,yq}=\beta_{yq}*(1[\text{year:quarter}=yq])*\text{RR-Depo}_{b,2007Q1} + \gamma\text{Post}_{yq}*\text{Bank Controls}_{b,2007Q1} + \delta_{f,b} + \delta_{f,yq} + \varepsilon_{f,b,yq}$$

where: Loan$_{f,b,yq}$ is is the log of total debt provided by bank b to firm f in year:quarter yq; 1[year:quarter=yq] is an indicator function with value 1 in year:quarter yq and 0 in other year:quarters; Bank Controls$_{b,2007Q1}$ is a vector of bank controls (as of 2007Q1) including: ROA, log Total Assets, Common Equity (rescaled by Total Assets), FX-Funds (rescaled by Total Assets), NPL (rescaled by Total Loans); $\delta_{f,b}$ is a vector of Firm*Bank fixed effects; $\delta_{f,yq}$ is a vector of Firm*Year:Quarter fixed effects and $\varepsilon_{f,b,yq}$ is an error term. Bank Controls are interacted with a post dummy, with value 1 (0) from 2007Q2 to 2008Q2 (from 2006Q1 to 2007Q1). The markers denote the point-estimate of the time-varying coefficients - representing the variation of loans relative to 2006Q1 induced by a unitary (100 p.p.) increase of RR-Depo$_{b,2007q1}$ - and the lines around them are 95% confidence interval. For reference, a 1 s.d. change in RR-Depo$_{b,2007q1}$ corresponds to 7.8 p.p.. Standard errors are double-clustered at the firm and bank*industry level.

Tables

TABLE 1**:** Summary Statistics

Panel A: Largest Sample for Regressions Exploiting Time Variation

VARIABLES	Definition: Timing	N	Mean	P25	P50	P75	SD
Loan-level Variables							
$Loan_{f,b,yq}$	Log(Loan): current year:quarter	1,475,369	16.843	15.317	17.051	18.507	2.573
$Provision_{f,b,yq-1}$	Loan Losses Provision (over Loan): 1Q-lagged	1,320,710	0.042	0.002	0.008	0.019	0.148
Macro Variables							
i_{yq-1}	Local Policy Rate: 1Q-lagged	1,475,369	0.075	0.062	0.073	0.092	0.014
$\Delta_{1y}GDP_{yq-1}$	1y-Growth of Local Policy Rate: 1Q-lagged	1,475,369	0.062	0.054	0.061	0.076	0.013
e_{yq-1}	Log(Exch. Rate: Pesos per 1 USD) : 1Q-lagged	1,475,369	7.692	7.608	7.724	7.754	0.090
CPI_{yq-1}	CPI (base: 2005Q1): 1Q-lagged	1,475,369	1.077	1.041	1.067	1.117	0.049
$MPspread^{US}_{yq-1}$	Local – US Policy Rate: 1Q-lagged	1,475,369	0.030	0.017	0.028	0.041	0.015
$SOVspread^{US}_{yq-1}$	Local – US (3-m) Sovereign yield: 1Q-lagged	1,475,369	0.036	0.023	0.033	0.047	0.018
CIP_{yq-1}	Deviations from CIP: 1Q-lagged	1,475,369	0.006	0.003	0.007	0.010	0.006
$\Delta_{2q}i_{yq-1}$	2q-Growth of Local Policy Rate: 1Q-lagged	1,475,369	0.005	-0.003	0.007	0.012	0.007
$\Delta_{1y}i_{yq-1}$	1y-Growth of Local Policy Rate: 1Q-lagged	1,475,369	0.008	-0.005	0.016	0.020	0.013
u_{yq-1}	Taylor Residuals (Rule 1) : 1Q-lagged	1,475,369	0.003	-0.002	0.006	0.006	0.006
u^{a}_{yq-1}	Taylor Residuals (Rule 2) : 1Q-lagged	1,475,369	0.002	-0.002	0.003	0.005	0.006
Bank-level Variables							
$FX\text{-}Funds_{b,yq-1}$	FX-Funds (over TA): 1Q-lagged	1,475,369	0.047	0.034	0.047	0.064	0.026
$SavingD_{b,yq-1}$	Savings Deposits (over TA): 1Q-lagged	1,475,369	0.353	0.303	0.348	0.400	0.073
$CheckingD_{b,yq-1}$	Checking Deposits (over TA): 1Q-lagged	1,475,369	0.137	0.108	0.126	0.173	0.045
$Size_{b,yq-1}$	Bank Log(TA): 1Q-lagged	1,475,369	30.301	30.02	30.383	30.704	0.523
$CET_{b,yq-1}$	Common Equity Capital (over TA): 1Q-lagged	1,475,369	0.042	0.032	0.039	0.050	0.013
$NPL_{b,yq-1}$	Non Perf. Loans (over Tot. Loans): 1Q-lagged	1,475,369	0.027	0.020	0.024	0.029	0.010
$ROA_{b,yq-1}$	Return on Assets: 1Q-lagged	1,475,369	0.014	0.008	0.012	0.019	0.007

This table shows summary statistics referred to the sample used in regressions for monetary policy rate which exploit time variation only, over the period 2005Q2 to 2008Q2. All the variables not defined as shares are expressed in (logs of) real Colombian Pesos with base year:quarter 2005Q1. In the definition of the macro variables, the Rule 1 is a Taylor Rule whereby the quarterly local policy rate is regressed against the (lagged) yearly inflation rate and the output gap; in Rule 2, against yearly inflation and log(GDP).

Panel B: Carry Trade Regressions

VARIABLES	Definition: Timing	N	Mean	P25	P50	P75	SD
Loan-level Variables							
$Loan_{f,b,yq}$	Log(Loan): current year:quarter	895,247	17.665	16.434	17.780	19.102	2.309
Macro Variables							
$MPspread^{US}_{yq-1}$	Local – US Policy Rate: 1Q-lagged	895,247	0.031	0.017	0.028	0.041	0.015
i_{yq-1}	Local Policy Rate: 1Q-lagged	895,247	0.075	0.062	0.073	0.092	0.014
$\Delta_{1y}GDP_{yq-1}$	1y-Growth of Local Policy Rate: 1Q-lagged	895,247	0.062	0.054	0.061	0.076	0.013
e_{yq-1}	Log(Exch. Rate: Pesos per 1 USD): 1Q-lagged	895,247	7.690	7.608	7.724	7.754	0.091
CPI_{yq-1}	CPI (base: 2005Q1): 1Q-lagged	895,247	1.078	1.041	1.067	1.117	0.050
VIX_{yq-1}	$Log(VIX)_{yq-1}$: 1Q-lagged	895,247	2.705	2.483	2.640	2.890	0.267
Oil_{yq-1}	Log(Brent $Price)_{yq-1}$: 1Q-lagged	895,247	4.286	4.090	4.246	4.488	0.239
$SOVspread^{US}_{yq-1}$	Local – US (3-month) Sovereign Yield: 1Q-lagged	895,247	0.036	0.023	0.033	0.047	0.018
CIP_{yq-1}	Deviations from CIP: 1Q-lagged	895,247	0.006	0.003	0.007	0.010	0.006
FP_{yq-1}	3-month COP-US$ Forward Premium: 1Q-lagged	895,247	0.030	0.020	0.027	0.040	0.017
Bank-level Variables							
$FX\text{-}Funds_{b,yq-1}$	FX-Funds (over TA): 1Q-lagged	895,247	0.046	0.030	0.047	0.063	0.026
$SavingD_{b,yq-1}$	Saving Deposits (over TA): 1Q-lagged	895,247	0.351	0.299	0.348	0.400	0.077
$CheckingD_{b,yq-1}$	Checking Deposits (over TA): 1Q-lagged	895,247	0.136	0.106	0.125	0.173	0.047
$Size_{b,yq-1}$	Bank Log(TA) : 1q-lagged	895,247	30.262	29.931	30.327	30.640	0.541
$CET_{b,yq-1}$	Common Equity Capital (over TA): 1Q-lagged	895,247	0.043	0.032	0.041	0.052	0.013
$NPL_{b,yq-1}$	Non Perf. Loans (over Tot. Loans): 1Q-lagged	895,247	0.027	0.020	0.024	0.030	0.011
$ROA_{b,yq-1}$	Return on Assets: 1Q-lagged	895,247	0.014	0.008	0.012	0.019	0.007
Firm-level Variables							
Firm $Risk_{f,pre}$	Mean Interest Payments (over Loan): 2005Q1-2007Q1	887,273	0.142	0.110	0.140	0.171	0.047
Short-Term $Debt_{f,pre}$	Mean Share of ST Debt: 2005Q1-2007Q1	887,530	0.341	0.080	0.274	0.548	0.296
$Default_{f,pre}$	At least 1 loan default: 2005Q1-2007Q1	887,874	0.305	0.000	0.000	1.000	0.460
$Supervised_{f,pre}$	Balance Sheet Supervised: 2005Q1- 2007Q1	895,247	0.302	0.000	0.000	1.000	0.459

This table shows summary statistics for the regression sample used for carry trade regressions, over the period 2005Q2 to 2008Q2. All the variables not defined as shares are expressed in (logs of) real Colombian Pesos with base year:quarter 2005Q1. In the definitions of bank variables, TA denotes banks total assets. In the definition of firm-level variables, ST Debt stands for Short-Term Debt, i.e. with maturity no longer than one year. $Default_{f,pre}$ is a 0/1 dummy. A loan default refers to a loan with payments which are at least 30 days past due. $Supervised_{f,pre}$ is a 0/1 dummy, with value 1 if the balance sheet is publicly supervised.

Panel C: Reserve-Requirements Regressions

VARIABLES	Definition: Timing	(1) N	(2) Mean	(3) P25	(4) P50	(5) P75	(6) SD
Loan-level Variables							
$Loan_{f,b,yq}$	Log(Loan): current year:quarter	742,950	17.658	16.437	17.778	19.096	2.314
$Provision_{f,b,yq-1}$	Loan Losses Provision (over Loan): 1Q-lagged	678,483	0.037	0.005	0.009	0.022	0.129
Bank-level Variables							
$RR\text{-}Depo_{b,2007Q1}$	Checking + Saving Dep. (over TA):2007Q1	742,950	0.514	0.483	0.534	0.574	0.078
$SavingD_{b,2007Q1}$	Saving Deposits (over TA): 2007Q1	742,950	0.381	0.309	0.392	0.400	0.071
$CheckingD_{b,2007Q1}$	Checking Deposits (over TA): 2007Q1	742,950	0.133	0.107	0.142	0.173	0.041
$Size_{b,2007Q1}$	Bank Log(TA) – 2007Q1	742,950	30.321	30.067	30.330	30.594	0.512
$CET_{b,2007Q1}$	Common Equity Capital (over TA): 2007Q1	742,950	0.040	0.032	0.034	0.050	0.013
$NPL_{b,2007Q1}$	Non Perf. Loans (over Tot. Loans): 2007Q1	742,950	0.025	0.020	0.023	0.024	0.007
$FX\text{-}Funds_{b,2007Q1}$	FX-Funds (over TA): 2007Q1	742,950	0.052	0.043	0.050	0.067	0.025
$ROA_{b,2007Q1}$	Return on Assets: 2007Q1	742,950	0.006	0.005	0.007	0.007	0.002
Firm-level Variables							
$Firm\ Risk_{f,pre}$	Mean Int. Paym. (over Loan): 2005Q1-2007Q1	734,976	0.142	0.110	0.141	0.173	0.047
$Short\text{-}Term\ Debt_{f,pre}$	Mean Share of ST Debt: 2005Q1-2007Q1	735,233	0.343	0.080	0.277	0.552	0.298
$Default_{f,pre}$	At least 1 loan default: 2005Q1-2007Q1	735,577	0.291	0	0	1	0.454
$Supervised_{f,pre}$	Balance Sheet Supervised: 2005Q1-2007Q1	742,950	0.294	0	0	1	0.455

This table shows summary statistics for the sample used in the regressions on Reserve Requirements policy – computed over the period 2006q1-2008q2. All the variables not defined as shares are expressed in (logs of) real Colombian Pesos with base year:quarter 2005Q1. In the definitions of bank variables, TA denotes banks total assets. In the definition of firm-level variables, ST Debt stands for Short-Term Debt, i.e. with maturity no longer than one year. $Default_{f,pre}$ is a 0/1 dummy: in its definition, a loan default refers to a loan with payments which are at least 30 days past due. $Supervised_{f,pre}$ is a 0/1 dummy, with value 1 if the balance sheet is publicly supervised.

TABLE 2: Local Monetary Policy Rate and Bank Credit

VARIABLES	(1)	(2)	(3) $Loan_{f,b,yq}$	(4)	(5)
$Post_{yq}*i_{yq-1}$	-3.452***	-5.644***	-5.382***	-6.321***	-3.586***
	(0.552)	(0.757)	(0.751)	(0.663)	(0.813)
i_{yq-1}	2.881***	4.333***	4.519***	4.688***	3.502***
	(0.433)	(0.566)	(0.562)	(0.499)	(0.517)
Observations	1,475,369	1,475,369	1,475,369	1,475,369	1,475,369
R-squared	0.674	0.674	0.678	0.832	0.832
Firm FE	Yes	Yes	Yes	-	-
Macro Control*Post	No	Yes	Yes	Yes	Yes
Bank FE	No	No	Yes	-	-
Firm*Bank FE	No	No	No	Yes	Yes
Bank Controls	No	No	No	No	Yes

This table shows the relation between bank credit and the local monetary policy rate. $Loan_{f,b,yq}$ is the log of total debt provided by bank b to firm f in year:quarter yq. i_{yq-1} is the lagged (by one quarter) local monetary policy rate. Macro controls include the lagged values of annual GDP growth, of the CPI index and of the (log) Peso-US$ exchange rate. Bank Controls include lagged (by one quarter): ROA, log Total Assets, Common Equity (rescaled by Total Assets), FX-Funds (rescaled by Total Assets), Savings Deposits (rescaled by Total Assets), Checking Deposits (rescaled by Total Assets). Standard errors are double-clustered at the Bank*Industry and at the Firm level. *** p<0.01, ** p<0.05, * p<0.1.

TABLE 3: Local Monetary Policy Rate and Bank Credit – Robustness

Panel A: Alternative Models

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
					$Loan_{f,b,yq}$			
$Post_{yq}$*i_{yq-1}	-10.775***	-3.430**	-2.414***	-8.806***	-3.081**	-3.586*	-3.586***	-3.586*
	(1.499)	(1.414)	(0.692)	(1.280)	(1.225)	(1.973)	(0.798)	(1.731)
i_{yq-1}	8.027***	0.242	3.301***	6.575***	0.505	3.502***	3.502***	3.502**
	(1.495)	(1.391)	(0.452)	(1.266)	(1.199)	(0.745)	(0.815)	(1.437)
Observations	1,362,608	1,203,805	1,475,369	1,362,608	1,203,805	1,475,369	1,475,369	1,475,369
R-squared	0.842	0.853	0.844	0.851	0.861	0.832	0.832	0.832
Loan-Size Weighted	No	No	Yes	Yes	Yes	No	No	No
Macro Control*Post	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm*Bank FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Bank Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Provision	No	Yes	No	No	Yes	No	No	No
Firm*Quarter FE	Yes	Yes	No	Yes	Yes	No	No	No

This table shows robustness exercises about the relation between bank credit and the local monetary policy rate. In all columns, the dependent variable is $Loan_{f,b,yq}$, the log of total debt provided by bank b to firm f in year:quarter yq. i_{yq-1} is the lagged (by one quarter) local monetary policy rate. In columns 1 through 2, we augment the baseline model with further controls and/or fixed effects. In columns 3 through 5, the model is estimated weighting variables by log-loan size. In columns 6 through 8, we apply alternative standard errors' clustering strategies. Macro controls include the lagged values of annual GDP growth, of the CPI index and of the (log) Peso-US$ exchange rate. Bank Controls include lagged (by one quarter): ROA, log Total Assets, Common Equity (rescaled by Total Assets), FX-Funds (rescaled by Total Assets), Savings Deposits (rescaled by Total Assets), Checking Deposits (rescaled by Total Assets). Provision is the lagged loan-level provision, rescaled by the loan value. Standard errors are double-clustered at the level of: Bank*Industry and Firm in columns 1 through 5; Bank and Firm in column 6; Bank, Firm and Year:Quarter in column 7; Bank*Industry, Firm and Year:Quarter in column 8. *** p<0.01, ** p<0.05, * p<0.1.

Panel B: Alternative Proxies of Local Monetary Policy Rate and Taylor Residuals

VARIABLES	(1)	(2)	(3) $Loan_{f,b,yq}$	(4)	(5)
$Post_{yq}$*$Proxy_{yq-1}$	-3.224***	-3.438**	-3.514***	-2.869***	-3.957***
	(0.737)	(1.546)	(1.068)	(0.792)	(0.742)
$Proxy_{yq-1}$	2.656***	3.696***	3.139***	2.047***	2.929***
	(0.353)	(0.469)	(0.377)	(0.553)	(0.495)
Observations	1,475,369	1,475,369	1,475,369	1,475,369	1,475,369
R-squared	0.832	0.832	0.832	0.832	0.832
$Proxy_{yq-1}$	$MPspread^{US}_{yq-1}$	$\Delta_{2q}i_{yq-1}$	$\Delta_{4q}i_{yq-1}$	u^1_{yq-1}	u^2_{yq-1}
Macro Controls*Post	Yes	Yes	Yes	Yes	Yes
Firm*Bank FE	Yes	Yes	Yes	Yes	Yes
Bank Controls	Yes	Yes	Yes	Yes	Yes

This table shows the relation between bank credit and different proxies of local monetary policy rate. In all columns, the dependent variable is $Loan_{f,b,yq}$, the log of total debt provided by bank b to firm f in year:quarter yq. Across different columns, we use alternative proxy of the monetary policy rate. In particular, in column 1, $MPspread^{US}_{yq-1}$ is the lagged spread between Colombian MP rate and the US Effective Federal Funds Rate. In column 2, $\Delta_{2q}i_{yq-1}$ is the lagged 2-quarter (half year) growth of Colombian MP rate. In column 3, $\Delta_{4q}i_{yq-1}$ is the lagged 1-year growth of the Colombian MP rate. In columns 4 and 5, respectively, u^1_{yq-1} and u^2_{yq-1} are Taylor residuals obtained from different policy rules. Macro controls include the lagged values of annual GDP growth, of the CPI index and of the (log) Peso-US$ exchange rate. Bank Controls include lagged (by one quarter): ROA, log Total Assets, Common Equity (rescaled by Total Assets), FX-Funds (rescaled by Total Assets), Savings Deposits (rescaled by Total Assets), Checking Deposits (rescaled by Total Assets). Standard errors are double-clustered at the Bank*Industry and at the Firm level. *** p<0.01, ** p<0.05, * p<0.1.

TABLE 4: Local Monetary Policy Rate and Bank Credit – Carry Trade Mechanism: Baseline Results

	(1)	(2)	(3)	(4)	(5)	(6)
			$Loan_{f,b,yq}$			
$MPspread^{US}_{yq-1}$*$FX\text{-}Funds_{b,yq-1}$*$Post_{yq}$		-109.378***	-105.444***	-109.090***	-109.358***	-280.971***
		(36.838)	(36.668)	(33.361)	(37.228)	(59.067)
$MPspread^{US}_{yq-1}$*$FX\text{-}Funds_{b,yq-1}$		55.557***	50.837**	53.148***	81.225***	144.609***
		(21.099)	(21.018)	(17.790)	(21.185)	(28.647)
$FX\text{-}Funds_{b,yq-1}$*$Post_{yq}$		-28.769	-35.367	-41.893	-9.303	19.526
		(32.990)	(32.976)	(30.034)	(35.459)	(36.522)
$FX\text{-}Funds_{b,yq-1}$	-0.950***	-43.477***	-36.322***	-35.900***	-51.056***	-58.943***
	(0.165)	(13.789)	(13.718)	(12.604)	(14.946)	(15.420)
$MPspread^{US}_{yq-1}$*$Post_{yq}$	-1.835*	-21.826***	-	-	-	-
	(1.014)	(2.748)				
$MPspread^{US}_{yq-1}$	2.069***	-0.212	-	-	-	-
	(0.507)	(0.880)				
Observations	895,247	895,247	895,247	895,247	895,247	895,247
R-squared	0.808	0.808	0.808	0.810	0.886	0.886
Firm*Bank FE	Yes	Yes	Yes	Yes	Yes	Yes
Macro Controls*Post	Yes	Yes	-	-	-	-
Bank Controls	Yes	-	-	-	-	-
Bank Controls*Post	No	Yes	Yes	Yes	Yes	-
FX-Funds*Macro Controls*Post	No	Yes	Yes	Yes	Yes	Yes
Year:Quarter FE	No	No	Yes	-	-	-
Industry*Year:Quarter FE	No	No	No	Yes	-	-
Firm*Year:Quarter FE	No	No	No	No	Yes	Yes
Bank Controls*$MPspread^{US}_{yq-1}$*Post	No	No	No	No	No	Yes

This table shows how carry trade strategies by local banks impacts the reaction of bank credit to the local monetary policy rate. The dependent variable, $Loan_{f,b,yq}$, is the log of total debt provided by bank b to firm f in year:quarter yq. $MPspread^{US}_{yq-1}$ is the difference between the (lagged) local monetary policy and the US Effective Federal Funds Rate. $FX\text{-}Funds_{b,yq-1}$ represents (lagged) bank FX-Funds (over Total Assets). Macro controls include the lagged values of annual GDP growth, of the CPI index and of the (log) Peso-US$ exchange rate. Bank Controls include lagged (by one quarter): ROA, log Total Assets, Common Equity (rescaled by Total Assets), Savings Deposits (rescaled by Total Assets) and Checking Deposits (rescaled by Total Assets). The sample consists of companies that borrowed from at least two banks. Standard errors are double-clustered at the Bank*Industry and at the Firm level. *** p<0.01, ** p<0.05, * p<0.1.

TABLE 5: Local Monetary Policy Rate and Bank Credit – Carry Trade Mechanism: Robustness

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
					$Loan_{f,b,yq}$				
$MPspread^{US}_{yq-1}$*$FX\text{-}Funds_{b,yq-1}$*$Post_{yq}$	-288.329***	-231.786***	-244.143***	-414.111***	-328.787***	-297.866***	-280.971**	-280.971**	-280.971**
	(56.723)	(51.920)	(50.517)	(45.070)	(37.933)	(36.776)	(112.997)	(126.842)	(112.734)
$MPspread^{US}_{yq-1}$*$FX\text{-}Funds_{b,yq-1}$	156.579***	136.535***	148.287***	157.006***	146.446***	159.482***	144.609**	144.609**	144.609***
	(27.880)	(25.537)	(25.089)	(21.804)	(18.954)	(18.602)	(50.188)	(49.075)	(46.711)
$FX\text{-}Funds_{b,yq-1}$*$Post_{yq}$	62.843*	38.665	73.801**	-86.157***	-43.921**	10.150	19.526	19.526	19.526
	(35.608)	(31.420)	(30.929)	(28.092)	(22.269)	(22.102)	(48.747)	(24.538)	(37.841)
$FX\text{-}Funds_{b,yq-1}$	-69.324***	-58.978***	-68.475***	-40.601***	-40.857***	-45.649***	-58.943**	-58.943	-58.943
	(15.308)	(13.223)	(13.205)	(9.791)	(8.217)	(8.344)	(26.579)	(34.548)	(37.108)
Observations	791,322	895,247	791,322	1,475,262	1,475,262	1,302,847	895,247	895,247	895,247
R-squared	0.894	0.889	0.898	0.834	0.846	0.857	0.886	0.886	0.886
Loan-Size Weighted	No	Yes	Yes	No	Yes	Yes	No	No	No
Firm*Bank FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
FX-Funds*Macro Controls*Post	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry*Year:Quarter FE	-	-	-	Yes	Yes	Yes	-	-	-
Firm*Year:Quarter FE	Yes	Yes	Yes	No	No	No	Yes	Yes	Yes
Bank Controls*$MPspread^{US}_{yq-1}$*Post	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Provision*$MPspread^{US}_{yq-1}$*Post	Yes	No	Yes	No	No	Yes	No	No	No
Companies	Multi-Bank	Multi-Bank	Multi-Bank	All	All	All	Multi-Bank	Multi-Bank	Multi-Bank

This table shows robustness exercises about how carry trade strategies by local banks impacts the reaction of bank credit to the local monetary policy rate. The dependent variable, $Loan_{f,b,yq}$, is the log of total debt provided by bank b to firm f in year:quarter yq. In columns 1 through 3, we augment the baseline model with either further controls and/or fixed effects, under OLS and WLS estimation. In columns 4 through 6, different versions of the model are estimated over a sample of companies consisting also of firms borrowing from bank only, whereas in the other columns Firm*Year:Quarter FE restrict the estimation sample to just those firms with at least two lenders (Multi-Bank firms). In columns 7 through 9, we apply alternative standard errors' clustering strategies. $MPspread^{US}_{yq-1}$ is the difference between the (lagged) local monetary policy and the US Effective Federal Funds Rate. $FX\text{-}Funds_{b,yq-1}$ represents (lagged) bank FX-Funds (over Total Assets). Macro controls include the lagged values of annual GDP growth, of the CPI index and of the (log) Peso-US$ exchange rate. Bank Controls include lagged (by one quarter): ROA, log Total Assets, Common Equity (rescaled by Total Assets), Savings Deposits (rescaled by Total Assets) and Checking Deposits (over Total Assets). Provision is the lagged loan-level provision, rescaled by the loan value. Standard errors are double-clustered at the level of: Bank*Industry and Firm in columns 1 through 6; Bank and Firm in column 7; Bank, Firm and Year:Quarter in column 8; Bank*Industry, Firm and Year:Quarter in column 9. *** p<0.01, ** p<0.05, * p<0.1.

TABLE 6: Local Monetary Policy Rate, Global Macroeconomic Factors and Bank Credit

	(1)	(2)	(3)
VARIABLES		$Loan_{f,b,yq}$	
$MPspread^{US}_{yq-1}$*$FX\text{-}Funds_{b,yq-1}$*$Post_{yq}$	-280.971***	-347.229***	-339.030**
	(59.067)	(92.252)	(145.969)
$MPspread^{US}_{yq-1}$*$FX\text{-}Funds_{b,yq-1}$	144.609***	166.740***	294.263***
	(28.647)	(30.001)	(37.001)
e_{yq-1}* $FX\text{-}Funds_{b,yq-1}$*$Post_{yq}$	3.025		
	(2.651)		
e_{yq-1}* $FX\text{-}Funds_{b,yq-1}$	-4.190**		
	(1.658)		
VIX_{yq-1}* $FX\text{-}Funds_{b,yq-1}$*$Post_{yq}$		5.868	
		(4.661)	
VIX_{yq-1}*$FX\text{-}Funds_{b,yq-1}$		-3.476**	
		(1.456)	
Oil_{yq-1}*$FX\text{-}Funds_{b,yq-1}$*$Post_{yq}$			-12.577
			(7.905)
Oil_{yq-1}*$FX\text{-}Funds_{b,yq-1}$			8.048***
			(1.185)
$FX\text{-}Funds_{b,yq-1}$*$Post_{yq}$	19.526	34.926	127.721***
	(36.522)	(21.981)	(49.459)
$FX\text{-}Funds_{b,yq-1}$	-58.943***	-91.424***	-157.988***
	(15.420)	(12.862)	(15.510)
Observations	895,247	895,247	895,247
R-squared	0.886	0.886	0.886
Firm*Bank FE	Yes	Yes	Yes
FX-Funds*Macro Controls*Post	Yes	Yes	Yes
Firm*Year:Quarter FE	Yes	Yes	Yes
Bank Controls*$MPspread^{US}_{yq-1}$*Post	Yes	Yes	Yes
H_0: e_{yq-1}*$FX\text{-}Funds_{b,yq-1}$+e_{yq-1}*$FX\text{-}Funds_{b,yq-1}$*$Post_{yq}$=0	0.58	-	-
H_0: e_{yq-1}*$FX\text{-}Funds_{b,yq-1}$=e_{yq-1}*$FX\text{-}Funds_{b,yq-1}$*$Post_{yq}$	0.06	-	-
H_0: VIX_{yq-1}*$FX\text{-}Funds_{b,yq-1}$+VIX_{yq-1}*$FX\text{-}Funds_{b,yq-1}$*$Post_{yq}$=0	-	0.58	-
H_0: VIX_{yq-1}*$FX\text{-}Funds_{b,yq-1}$=VIX_{yq-1}*$FX\text{-}Funds_{b,yq-1}$*$Post_{yq}$	-	0.08	-
H_0: Oil_{yq-1}*$FX\text{-}Funds_{b,yq-1}$+Oil_{yq-1}*$FX\text{-}Funds_{b,yq-1}$*$Post_{yq}$=0	-	-	0.56
H_0: Oil_{yq-1}*$FX\text{-}Funds_{b,yq-1}$=Oil_{yq-1}*$FX\text{-}Funds_{b,yq-1}$*$Post_{yq}$	-	-	0.01

This table shows how bank FX-funding influences bank credit reaction to global macroeconomic and external factors. The dependent variable, $Loan_{f,b,yq}$, is the log of total debt provided by bank b to firm f in year:quarter yq. $MPspread^{US}_{yq-1}$ is the difference between the (lagged) local monetary policy and the FED Effective Funds Rate. e_{yq-1} is the lagged (log) nominal exchange rate, expressed as pesos per 1 USD, so that an increase denotes a depreciation of the Colombian peso against the USD. VIX_{yq-1} is the lagged (log) VIX index, whereas Oil_{yq-1} is the lagged (log) Brent oil price. $FX\text{-}Funds_{b,yq-1}$ represents (lagged) bank FX-Funds (over Total Assets). Macro controls include the lagged values of annual GDP growth, of the CPI index and of the (log) Peso-US$ exchange rate. Bank Controls include lagged (by one quarter): ROA, log Total Assets, Common Equity (rescaled by Total Assets), Savings Deposits (rescaled by Total Assets) and Checking Deposits (over Total Assets). The last six rows report the p-values for the tests with null hypothesis specified in the first column. Standard errors are double-clustered at the Bank*Industry and Firm level. *** p<0.01, ** p<0.05, * p<0.1.

TABLE 7: Local Monetary Policy Rate and Bank Credit – Carry Trade Mechanism– *Breakdown by Currency* (smaller sample)

VARIABLES	(1) Peso $Loan_{f,b,yq}$	(2) FX $Loan_{f,b,yq}$	(3) (Peso Loan/ $Loan)_{f,b,yq}$	(4) $Loan_{f,b,yq}$
$MPspread^{US}_{yq-1}$*$FX\text{-}Funds_{b,yq-1}$*$Post_{yq}$	-417.566***	230.785	-19.692**	-202.724**
	(113.575)	(427.576)	(8.780)	(96.183)
$MPspread^{US}_{yq-1}$*$FX\text{-}Funds_{b,yq-1}$	222.817***	87.433	6.084	169.646***
	(54.926)	(157.764)	(4.440)	(46.026)
$FX\text{-}Funds_{b,yq-1}$*$Post_{yq}$	15.964	-430.059	-2.812	96.400^
	(75.847)	(325.617)	(7.105)	(63.617)
$FX\text{-}Funds_{b,yq-1}$	-149.157***	14.977	-14.486***	-97.449***
	(30.872)	(119.729)	(2.915)	(26.425)
Observations	315,692	22,686	322,775	322,775
R-squared	0.835	0.891	0.785	0.857
Firm*Bank FE	Yes	Yes	Yes	Yes
FX-Funds*Macro Controls*Post	Yes	Yes	Yes	Yes
Firm*Year:Quarter FE	Yes	Yes	Yes	Yes
Bank Controls*$MPspread^{US}_{yq-1}$*Post	Yes	Yes	Yes	Yes

This table shows how carry trade strategies by local banks impacts the reaction of bank credit to local monetary policy, depending on the currency of denomination of the loans. In column 1, the dependent variable is the log of peso loans provided by bank b to firm f in Year:Quarter yq. Peso $Loan_{f,b,yq}$ is the log of total peso-denominated debt provided by bank b to firm f in year:quarter yq. FX $Loan_{f,b,yq}$ is the log of total FX-denominated debt provided by bank b to firm f in year:quarter yq. (Peso Loan/ $Loan)_{f,b,yq}$ represents the share of peso-denominated debt out of the total debt provided by bank b to firm f in year:quarter yq. $Loan_{f,b,yq}$, is the log of total debt provided by bank b to firm f in year:quarter yq. $MPspread^{US}_{yq-1}$ is the difference between the (lagged) local monetary policy and the FED Effective Funds Rate. $FX\text{-}Funds_{b,yq-1}$ represents (lagged) bank FX-Funds (over Total Assets). Macro controls include the lagged values of annual GDP growth, of the CPI index and of the (log) Peso-US$ exchange rate. Bank Controls include lagged (by one quarter): ROA, log Total Assets, Common Equity (rescaled by Total Assets), Savings Deposits (rescaled by Total Assets), Checking Deposits (over Total Assets). Standard errors are double-clustered at the Bank*Industry level and at the Firm level. *** p<0.01, ** p<0.05, * p<0.1.

TABLE 8: Local Monetary Policy Rate and Bank Credit – Deviations from CIP and Carry Trade Mechanism

VARIABLES	(1)	(2) Loan$_{f,b,yq}$	(3)	(4) Peso Loan$_{f,b,yq}$	(5) FX Loan$_{f,b,yq}$	(6) (Peso Loan/ Loan)$_{f,b,yq}$
FX-Funds$_{b,yq-1}$*SOVspread$^{US}_{yq}$*Post$_{yq}$	-184.667***					
	(50.870)					
FX-Funds$_{b,yq-1}$*CIP$_{yq-1}$*Post$_{yq}$		-246.815***	-430.574***	-851.442***	-361.149	-69.018***
		(66.541)	(81.650)	(172.725)	(627.477)	(14.050)
FX-Funds$_{b,yq-1}$*FP$_{yq-1}$*Post$_{yq}$			-116.129***	-163.204*	489.968^	-8.330
			(44.903)	(88.504)	(331.862)	(6.612)
FX-Funds$_{b,yq-1}$*SOVspread$^{US}_{yq}$	78.632***					
	(21.936)					
FX-Funds$_{b,yq-1}$*CIP$_{yq-1}$		76.099***	132.700***	141.584*	6.630	1.915
		(26.386)	(45.529)	(85.418)	(257.284)	(7.084)
FX-Funds$_{b,yq-1}$*FP$_{yq-1}$			64.305**	94.742*	52.404	2.562
			(25.905)	141.584*	(135.489)	(4.090)
FX-Funds$_{b,yq-1}$*Post$_{yq}$	-48.473	-22.487	87.595**	180.273**	45.107	23.786***
	(46.225)	(57.780)	(39.405)	(79.683)	(295.494)	(6.734)
FX-Funds$_{b,yq-1}$	-0.741	-10.806	-4.120	-28.259	44.689	-5.787**
	(12.038)	(12.406)	(13.070)	(25.262)	(99.145)	(2.251)
Observations	895,247	895,247	895,247	315,692	22,686	322,775
R-squared	0.886	0.886	0.886	0.835	0.891	0.785
Firm*Bank FE	Yes	Yes	Yes	Yes	Yes	Yes
FX-Funds*Macro Controls*Post	Yes	Yes	Yes	Yes	Yes	Yes
Firm*Year:Quarter FE	Yes	Yes	Yes	Yes	Yes	Yes
Bank Controls*Int Rate*Post	Yes	Yes	Yes	Yes	Yes	Yes

This table shows how carry trade strategies by local banks impacts the reaction of bank credit to local monetary policy depending on deviations from CIP. Loan$_{f,b,yq}$ is the log of total debt provided by bank b to firm f in year:quarter yq. SOVspread$^{US}_{yq-1}$ is the difference between the (lagged) yields on 3-month Colombian and US Sovereign bonds. CIP$_{yq-1}$ is the (lagged) deviation from CIP based on the 3-month yield-differential between Colombian and US Sovereign bonds, computed by Du and Schreger (2016). FP$_{yq-1}$ is the 3-month Peso/US$ forward premium, expressing the difference between the latter two variables, i.e. the 3-month Colombia-US Sovereign yield differential not imputable to deviations from CIP. FX-Funds$_{b,yq-1}$ represents (lagged) bank FX-Funds (over Total Assets). Macro controls include the lagged values of annual GDP growth, of the CPI index and of the (log) Peso-US$ exchange rate. Bank Controls include lagged (by one quarter): ROA, log Total Assets, Common Equity (rescaled by Total Assets), Savings Deposits (rescaled by Total Assets), Checking Deposits (rescaled by Total Assets). Standard errors are double-clustered at the Bank*Industry and at the Firm level. *** p<0.01, ** p<0.05, * p<0.1, ^ p<0.15.

TABLE 9: Local Monetary Policy Rate and Bank Credit – Carry Trade Mechanism: Firms Heterogeneity

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	Firm Risk$_{f,pre}$				Short TermDebt (maturity ≤ 1y)$_{f,pre}$				30-day Past Due$_{f,pre}$		Supervised$_{f,pre}$	
VARIABLES	Q=1	Q=2	Q=3	Q=4	Q=1	Q=2	Q=3	Q=4	No	Yes	Yes	No
$MPspread^{US}_{yq-1}$*FX-Funds$_{b,yq-1}$*Post$_{yq}$	-94.897	-189.267*	-375.565***	-577.257***	-39.874	-430.129***	-257.162**	-342.661***	-233.226***	-374.137***	-175.004	-337.660***
	(126.913)	(98.445)	(108.555)	(122.708)	(117.157)	(111.560)	(103.053)	(119.001)	(70.705)	(97.968)	(108.069)	(67.337)
$MPspread^{US}_{yq-1}$*FX-Funds$_{b,yq-}$	-31.842	152.189***	322.070***	279.582***	-14.109	168.322***	154.613***	240.740***	111.753***	194.653***	194.357***	124.372***
	(56.991)	(48.387)	(53.652)	(67.529)	(56.007)	(52.554)	(50.298)	(60.923)	(35.249)	(46.441)	(52.406)	(32.643)
FX-Funds$_{b,yq-1}$*Post$_{yq}$	-107.802	31.230	145.325**	37.102	51.154	-3.949	53.506	-16.872	17.720	20.539	138.991*	-36.626
	(76.460)	(64.840)	(71.109)	(86.591)	(76.382)	(67.801)	(67.132)	(79.625)	(45.564)	(63.616)	(73.828)	(40.956)
FX-Funds$_{b,yq-1}$	-20.129	-77.012***	-99.150***	-68.918	6.903	-91.306***	-67.837**	-67.443**	-67.479***	-42.008*	-126.701***	-27.548
	(31.973)	(26.460)	(30.018)	(43.452)	(30.011)	(28.247)	(28.826)	(34.031)	(19.552)	(24.355)	(29.847)	(17.506)
Observations	228,530	254,094	224,103	180,546	192,636	226,841	241,978	225,818	617,034	270,840	270,253	624,994
R-squared	0.884	0.863	0.859	0.886	0.914	0.876	0.873	0.876	0.881	0.894	0.862	0.887
Firm*Bank FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
FX-Funds*Macro Controls*Post	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm*Year:Quarter FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Bank Controls*Int Rate*Post	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

This table shows how carry trade strategies by local banks impact the reaction of bank credit to local monetary policy, across different groups of companies. In columns 1 through 4, companies are sorted according to the distribution of the average interest payments over total assets paid between 2005Q1 and 2007Q1. In columns 5 through 8, companies are sorted according to the distribution of the average share of bank debt with maturity no longer than one year borrowed between 2005Q1 and 2007Q1. In columns 1 through 8, Q=j denotes that a company falls in the j-th quartile of the relevant distribution, j={1,2,3,4}. In columns 9 through 10, companies are divided depending on whether they are 30 days past due with respect to at least one bank loan between 2005Q1 and 2007Q1. Finally, in columns 11 and 12, companies are sorted according to whether their balance sheet is publicly supervised at least once between 2005Q1 and 2008Q2. Loan$_{f,b,yq}$ is the log of total debt provided by bank b to firm f in year:quarter yq. $MPspread^{US}_{yq-1}$ is the difference between the (lagged) local monetary policy and the FED Effective Funds Rate. FX-Funds$_{b,yq-1}$ represents (lagged) bank FX-Funds (over Total Assets). Macro controls include the lagged values of annual GDP growth, of the CPI index and of the (log) Peso-US$ exchange rate. Bank Controls include lagged (by one quarter): ROA, log Total Assets, Common Equity (rescaled by Total Assets), Savings Deposits (rescaled by Total Assets), Checking Deposits (rescaled by Total Assets). Standard errors are double-clustered at the Bank*Industry and at the Firm level. *** p<0.01, ** p<0.05, * p<0.1.

TABLE 10: Reserve Requirement Shock and Bank Credit

	(1)	(2)	(3)	(4)	(5)	(6)
			$Loan_{f,b,yq}$			
$RR\text{-}Depo_{b,2007Q1}$	0.817***	-	-	-	-	
	(0.295)					
$Post_{yq}$	-2.366***	-0.455	-	-	-	
	(0.577)	(0.528)				
$Post_{yq}*RR\text{-}Depo_{b,2007Q1}$	-1.542***	-0.994***	-1.017***	-1.048***	-0.697***	
	(0.196)	(0.181)	(0.181)	(0.160)	(0.178)	
$Post_{yq}*SavingD_{b,2007Q1}$						-0.419**
						(0.179)
$Post_{yq}*CheckingD_{b,2007Q1}$						-1.845***
						(0.281)
Observations	742,950	742,950	742,950	742,950	742,950	742,950
R-squared	0.536	0.829	0.829	0.830	0.897	0.897
Bank Controls*Post	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	-	-	-	-	-
Firm*Bank FE	No	Yes	Yes	Yes	Yes	Yes
Year:Quarter	No	No	Yes	-	-	-
Industry*Year:Quarter	No	No	No	Yes	-	-
Firm*Year:Quarter	No	No	No	No	Yes	Yes

This table shows the evolution of bank credit in reaction to the Reserve Requirement (RR) shock. The dependent variable is $Loan_{f,b,yq}$, i.e. the log of total debt provided by bank b to firm f in year:quarter yq. RR-Depo is the sum of savings ($SavingD_{b,2007Q1}$) and checking ($CheckingD_{b,2007Q1}$) deposits, both rescaled by total assets, as of 2007Q1. In Panel B, in all columns we include only firms that borrowed from at least two banks. Bank Controls is a vector of bank controls (as of 2007Q1) including: ROA, log Total Assets, Common Equity (rescaled by Total Assets), FX-Funds (rescaled by Total Assets), NPL (rescaled by Total Loans). The $Post_{yq}$ dummy has value 1 from 2007Q2 onward and value 0 before. Standard errors are double-clustered at the Bank*Industry and at the Firm level. *** p<0.01, ** p<0.05, * p<0.1.

TABLE 11: Reserve Requirement Shock and Bank Credit: Robustness

Panel A: Alternative Models

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
					$Loan_{f,b,yq}$				
$Post_{yq}$*RR-$Depo_{b,2007Q1}$	-0.675***	-0.647***	-0.612***	-0.844***	-0.871***	-0.994***	-0.697^	-0.697*	-0.697**
	(0.183)	(0.159)	(0.166)	(0.134)	(0.116)	(0.119)	(0.414)	(0.367)	(0.221)
Observations	640,136	742,950	640,136	1,219,366	1,219,366	1,049,099	742,950	742,950	742,950
R-squared	0.908	0.900	0.911	0.851	0.862	0.877	0.897	0.897	0.897
Loan-Size Weighted	No	Yes	Yes	No	Yes	Yes	No	No	No
Firm*Bank FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
FX-Funds*Macro Controls*Post	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry*Year:Quarter FE	-	-	-	Yes	Yes	Yes	-	-	-
Firm*Year:Quarter FE	Yes	Yes	Yes	No	No	No	Yes	Yes	Yes
Bank Controls*Int Rate*Post	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Provision*MP-spread*Post	Yes	No	Yes	No	No	Yes	No	No	No

This table shows robustness exercises about the reaction of bank credit to the Reserve Requirement (RR) shock. In columns 1 through 3, we augment the baseline model with either further controls and/or fixed effects, under OLS and WLS estimation. In columns 4 through 6, different versions of the model are estimated over a sample of companies consisting of also firms borrowing from bank only, whereas in the other columns Firm*Year:Quarter FE restrict the estimation sample to just those firms with at least two lenders (Multi-Bank firms). In columns 7 through 9, we apply alternative standard errors' clustering strategies. The dependent variable, $Loan_{f,b,yq}$, is the log of total debt provided by bank b to firm f in year:quarter yq. RR-Depo is the sum of savings ($SavingD_{b,2007Q1}$) and checking ($CheckingD_{b,2007Q1}$) deposits, both rescaled by total assets, as of 2007Q1. In Panel B, in all columns we include only firms borrowing from at least two banks. Bank Controls is a vector of bank controls (as of 2007Q1) including: ROA, log Total Assets, Common Equity (rescaled by Total Assets), FX-Funds (rescaled by Total Assets), NPL (rescaled by Total Loans). The $Post_{yq}$ dummy has value 1 from 2007Q2 onward and value 0 before. Provision is the lagged loan-level provision, rescaled by the loan value. Standard errors are double-clustered at the level of: Bank*Industry and Firm in columns 1 through 6; Bank and Firm in column 7; Bank, Firm and Year:Quarter in column 8; Bank*Industry, Firm and Year:Quarter in column 9. *** p<0.01, ** p<0.05, * p<0.1, ^ p<0.12.

Panel B: Cross-Sectional Regressions

VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	$\Delta_1 Loan_{fb,2007Q2}$		$\Delta_2 Loan_{fb,2007Q3}$		$\Delta_3 Loan_{fb,2007Q4}$		$\Delta_4 Loan_{fb,2008Q1}$		$\Delta_5 Loan_{fb,2008Q2}$	
$RR\text{-}Depo_{b,2007Q1}$	-0.285**		-0.673***		-1.042***		-0.778***		-0.842***	
	(0.134)		(0.169)		(0.207)		(0.219)		(0.242)	
$SavingD_{b,2007Q1}$		-0.181		-0.548***		-0.870***		-0.458**		-0.498**
		(0.133)		(0.169)		(0.209)		(0.222)		(0.246)
$CheckingD_{b,2007Q1}$		-0.738***		-1.222***		-1.794***		-2.158***		-2.302***
		(0.214)		(0.274)		(0.326)		(0.347)		(0.367)
Observations	66,758	66,758	63,993	63,993	60,865	60,865	58,921	58,921	57,199	57,199
R-squared	0.378	0.378	0.393	0.394	0.405	0.405	0.414	0.414	0.425	0.426
Bank Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

This table shows the evolution of bank credit in reaction to the Reserve Requirement (RR) shock. We perform cross-sectional regressions. The dependent variable is the difference between $Loan_{f,b,2007Q1+j}$ and $Loan_{f,b,2007Q1}$, signaled by the operator Δ_j, j={1,2,3,4,5}. Note that the starting year:quarter is always 2007Q1, the year:quarter before the RR-Shock. RR-Depo is the sum of savings ($SavingD_{b,2007Q1}$) and checking ($CheckingD_{b,2007Q1}$) deposits, both over total assets, as of 2007Q1. In Panel B, in all columns we include only firms borrowing from at least two banks. Bank Controls is a vector of bank controls (as of 2007Q1) including: ROA, log Total Assets, Common Equity (rescaled by Total Assets), FX-Funds (rescaled by Total Assets), NPL (rescaled by Total Loans). The $Post_{yq}$ dummy has value 1 from 2007Q2 onward and value 0 before. Standard errors are double-clustered at the Bank*Industry and at the Firm level. *** p<0.01, ** p<0.05, * p<0.1.

Panel C: Placebo test

	(1)	(2)
VARIABLES	$Loan_{f,b,yq}$	
$Post(Fake)_{yq}$*$RR\text{-}Depo_{b,2005Q4}$	0.368	
	(0.256)	
$Post(Fake)_{yq}$ *$SavingD_{b,2005Q4}$		0.800***
		(0.299)
$Post(Fake)_{yq}$ *$CheckingD_{b,2005Q4}$		0.313
		(0.255)
Observations	486,201	486,201
R-squared	0.903	0.903
Bank Controls*Post	Yes	Yes
Firm*Bank FE	Yes	Yes
Firm*Year:Quarter	Yes	Yes

This table performs a placebo test. The sample goes from 2005Q1 to 2006Q4. The dependent variable is $Loan_{f,b,yq}$, i.e. the log of total debt provided by bank b to firm f in year:quarter yq. Banks variables are measured at 2005Q4, a year:quarter with no RR-intervention. RR-Depo is the sum of savings ($SavingD_{b,2005Q4}$) and checking ($CheckingD_{b,2005Q4}$) deposits, both over total assets, as of 2005Q4. Bank Controls is a vector of bank controls (as of 2005Q4) including: ROA, log Total Assets, Common Equity and FX-Funds (both rescaled by Total Assets), NPL (rescaled by Total Loans). The $Post(Fake)_{yq}$ dummy has value 1 from 2006Q1 onward and 0 before. Standard errors are double-clustered at the Bank*Industry and at the Firm level. *** p<0.01, ** p<0.05, * p<0.1.

TABLE 12: Reserve Requirement Shock and Bank Credit – Firms Heterogeneity

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	Dependent Variable: $Loan_{f,b,yq}$											
	Firm $Risk_{f,pre}$				Short-Term Debt (maturity ≤ 1y)$_{f,pre}$				30-day Past $Due_{f,pre}$		$Supervised_f$	
	Q=1	Q=2	Q=3	Q=4	Q=1	Q=2	Q=3	Q=4	No	Yes	Yes	No
$Post_{yq}$*RR-$Depo_{b,2007Q1}$	0.290	-0.667**	-1.169***	-1.978***	-0.517	-0.375	-1.083***	-0.770**	-0.680***	-0.724**	-0.066	-0.994***
	(0.378)	(0.298)	(0.362)	(0.411)	(0.414)	(0.343)	(0.326)	(0.341)	(0.216)	(0.313)	(0.327)	(0.217)
Observations	190,069	211,012	184,976	148,919	160,095	190,709	198,522	185,650	521,480	214,097	218,269	524,681
R-squared	0.894	0.880	0.873	0.894	0.921	0.890	0.885	0.888	0.892	0.906	0.877	0.897
Bank Controls*Post	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm*Bank FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm*Year:Quarter	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

This table evaluates the effects of the Reserve Requirement on shock on bank credit, across different groups of companies In columns 1 through 4, companies are sorted according to the distribution of the average interest payments over total assets paid between 2005Q1 and 2007Q1. In columns 5 through 8, companies are sorted according to the distribution of the average share of bank debt with maturity no longer than one year borrowed between 2005Q1 and 2007Q1. In columns 1 through 8, Q=j denotes that a company falls in the j-th quartile of the relevant distribution, j={1,2,3,4}. In columns 9 through 10, companies are divided depending on whether they are 30 days past due for at least one bank loan between 2005Q1 and 2007Q1. Finally, in columns 11 and 12, companies are sorted according to whether their balance sheet is publicly supervised at least once between 2005Q1 and 2008Q2. $Loan_{f,b,yq}$ is the log of total debt provided by bank b to firm f in year:quarter yq. RR-Depo is the sum of savings ($SavingD_{b,2007Q1}$) and checking ($CheckingD_{b,2007Q1}$) deposits, both over total assets, as of 2007Q1. In Panel B, in all columns we include only firms that borrowed from at least two banks. Bank Controls is a vector of bank controls (as of 2007Q1) including: ROA, log Total Assets, Common Equity (rescaled by Total Assets), FX-Funds (rescaled by Total Assets), NPL (rescaled by Total Loans). The $Post_{yq}$ dummy has value 1 from 2007Q2 onward and value 0 before. In all columns, the regressions include companies that borrowed from at least two banks. Standard errors are double-clustered at the Bank*Industry and at the Firm level. *** p<0.01, ** p<0.05, * p<0.1.

TABLE 13: Capital Controls, Domestic Macroprudential Policy and the Bank Lending Channel of Monetary Policy

	(1)	(2)	(3)	(4)
		$Loan_{f,b,yq}$		
$MPspread^{US}_{yq-1}$*FX-$Funds_{b,yq-1}$*$Post_{yq}$	-109.378***	-105.444***	-109.090***	-109.358***
	(36.838)	(36.668)	(33.361)	(37.228)
$MPspread^{US}_{yq-1}$*FX-$Funds_{b,yq-1}$	55.557***	50.837**	53.148***	81.225***
	(21.099)	(21.018)	(17.790)	(21.185)
FX-$Funds_{b,yq-1}$*$Post_{yq}$	-28.769	-35.367	-41.893	-9.303
	(32.990)	(32.976)	(30.034)	(35.459)
FX-$Funds_{b,yq-1}$	-43.477***	-36.322***	-35.900***	-51.056***
	(13.789)	(13.718)	(12.604)	(14.946)
$CheckingD_{b,yq-1}$*$Post_{yq}$	-0.417***	-0.411***	-0.425***	-0.367**
	(0.142)	(0.143)	(0.143)	(0.158)
$CheckingD_{b,yq-1}$	0.003	-0.136	-0.133	-0.055
	(0.107)	(0.120)	(0.103)	(0.123)
$SavingD_{b,yq-1}$*$Post_{yq}$	-0.705***	-0.696***	-0.715***	-0.679***
	(0.110)	(0.113)	(0.098)	(0.108)
$SavingD_{b,yq-1}$	-0.066	-0.069	-0.075	-0.064
	(0.111)	(0.119)	(0.107)	(0.121)
Observations	895,247	895,247	895,247	895,247
R-squared	0.808	0.808	0.810	0.886
Firm*Bank FE	Yes	Yes	Yes	Yes
Macro Controls*Post	Yes	-	-	-
Bank Controls	-	-	-	-
Bank Controls*Post	Yes	Yes	Yes	Yes
FX-Funds*Macro Controls*Post	Yes	Yes	Yes	Yes
Year:Quarter FE	No	Yes	-	-
Industry*Year:Quarter FE	No	No	Yes	-
Firm*Year:Quarter FE	No	No	No	Yes

This table shows the impact of capital controls and reserve requirements on bank credit. The dependent variable, $Loan_{f,b,yq}$, is the log of total debt provided by bank b to firm f in year:quarter yq. $MPspread^{US}_{yq-1}$ is the difference between the (lagged) local monetary policy and the FED Effective Funds Rate. FX-$Funds_{b,yq-1}$ represents (lagged) bank FX-Funds (over Total Assets). $SavingD_{b,yq-1}$ denotes (lagged) bank savings deposits (over total assets). $CheckingD_{b,yq-1}$ denotes (lagged) bank checking deposits (over total assets). Macro controls include the lagged values of annual GDP growth, of the CPI index and of the (log) Peso-US$ exchange rate. Bank Controls include lagged (by one quarter): ROA, log Total Assets, Common Equity (rescaled by Total Assets). The sample consists only of companies that borrowed from at least two banks. Standard errors are double-clustered at the Bank*Industry and at the Firm level. *** p<0.01, ** p<0.05, * p<0.1.

Appendix

TABLE A1: Summary Statistics for Carry Trade Regressions (Smaller Sample with Currency Breakdown of Loan Volume)

VARIABLES	Definition: timing	(1) N	(2) Mean	(3) P25	(4) P50	(5) P75	(6) SD
Loan-level Variables							
Peso $\text{Loan}_{f,b,yq}$	Log(Peso Loan): current year:quarter	315,692	18.347	17.176	18.602	19.890	2.544
FX $\text{Loan}_{f,b,yq}$	Log(FX Loan): current year:quarter	22,686	19.525	18.570	19.797	21.002	2.359
$(\text{Peso Loan/ Loan})_{f,b,yq}$	Peso Loan / Total Loan: current year:quarter	322,775	0.941	1.000	1.000	1.000	0.210
Bank-level variables							
$\text{FX-Funds}_{b,yq-1}$	FX-Funds (over TA): 1Q-lagged	322,775	0.046	0.030	0.046	0.062	0.025
$\text{SavingD}_{b,yq-1}$	Savings Deposits (over TA): 1Q-lagged	322,775	0.346	0.287	0.341	0.400	0.079
$\text{CheckingD}_{b,yq-1}$	Checking Deposits (over TA): 1Q-lagged	322,775	0.135	0.105	0.123	0.173	0.047
$\text{Size}_{b,yq-1}$	Bank Log(TA) : 1Q-lagged	322,775	16.399	16.052	16.425	16.798	0.549
$\text{CET}_{b,yq-1}$	Common Equity Capital (over TA): 1Q-lagged	322,775	0.043	0.032	0.042	0.052	0.013
$\text{NPL}_{b,yq-1}$	Non Perf. Loans (over Tot. Loans): 1Q-lagged	322,775	0.027	0.020	0.024	0.029	0.010
$\text{ROA}_{b,yq-1}$	Return on Assets: 1Q-lagged	322,775	0.014	0.008	0.012	0.018	0.007

TABLE A2: Summary Statistics for Reserve-Requirements Policy Regressions (Placebo Sample in Table 10)

VARIABLES	Definition: timing	(1) N	(2) Mean	(3) P25	(4) P50	(5) P75	(6) SD
Loan-level Variables							
$Loan_{f,b,yq}$	Log(Loan): current year:quarter	486,201	17.638	16.386	17.745	19.088	7.846
Bank-level variables							
$RR\text{-}Depo_{b,2005Q4}$	Checking + Savings Dep. (over TA): 2005Q4	486,201	0.491	0.459	0.520	0.530	0.079
$SavingD_{b,2005Q4}$	Savings Deposits (over TA): 2005Q4	486,201	0.332	0.273	0.346	0.373	0.070
$CheckingD_{b,2005Q4}$	Checking Deposits (over TA): 2005Q4	486,201	0.159	0.128	0.145	0.226	0.058
$Size_{b,2005Q4}$	Bank Log(TA) – 2005Q4	486,201	16.339	16.092	16.375	16.598	0.523
$CET_{b,2005Q4}$	Common Equity Capital (over TA): 2005Q4	486,201	0.040	0.030	0.034	0.050	0.013
$NPL_{b,2005Q4}$	Non Perf. Loans (over Tot. Loans): 2005Q4	486,201	0.024	0.020	0.020	0.022	0.009
$FX\text{-}Funds_{b,2005Q4}$	FX-Funds (over TA): 2005Q4	486,201	0.045	0.028	0.038	0.059	0.029
$ROA_{b,2005Q4}$	Return on Assets: 2005Q4	486,201	0.024	0.016	0.028	0.031	0.007

TABLE A3: Local Monetary Policy, Foreign *vs* Domestic Bank Funding and Credit

	(1)
	$Loan_{f,b,yq}$
$MPspread^{US}_{yq-1}$*FX-$Funds_{b,yq-1}$*$Post_{yq}$	-280.971***
	(59.067)
$MPspread^{US}_{yq-1}$*FX-$Funds_{b,yq-1}$	144.609***
	(28.647)
$MPspread^{US}_{yq-1}$*$SavingD_{b,yq-1}$*$Post_{yq}$	-11.210
	(10.749)
$MPspread^{US}_{yq-1}$*$SavingD_{b,yq-1}$	-7.413
	(7.117)
$MPspread^{US}_{yq-1}$*$CheckingD_{b,yq-1}$*$Post_{yq}$	-13.325
	(9.307)
$MPspread^{US}_{yq-1}$*$CheckingD_{b,yq-1}$	8.429
	(14.103)
R-squared	0.886
Firm*Bank FE	Yes
FX-Funds*Macro Controls*Post	Yes
Firm*Year:Quarter FE	Yes
Bank Controls*$MPspread^{US}_{yq-1}$*Post	Yes
Bank Controls*i_{yq-1}*Post	No

This table shows the impact of the monetary policy rate spread on bank credit, conditional on different bank funding structures. (Note: this table reproduces column 6 of Table 4, displaying additional coefficients). The dependent variable, $Loan_{f,b,yq}$, is the log of total debt provided by bank b to firm f in year:quarter yq. $MPspread^{US}_{yq-1}$ is the difference between the (lagged) local monetary policy and the FED Effective Funds Rate. FX-$Funds_{b,yq-1}$ represents (lagged) bank FX-Funds (over Total Assets). $SavingD_{b,yq-1}$ denotes (lagged) bank savings deposits (over total assets). $CheckingD_{b,yq-1}$ denotes (lagged) bank checking deposits (over total assets). Macro controls include the lagged values of annual GDP growth, of the CPI index and of the (log) Peso-US$ exchange rate. Bank Controls include lagged (by one quarter): ROA, log Total Assets, Common Equity (rescaled by Total Assets). The sample includes only those companies that borrowed from at least two banks. Standard errors are double-clustered at the Bank*Industry and at the Firm level. *** p<0.01, ** p<0.05, * p<0.1.

TABLE A4: Using the Monetary Policy Rate instead of the Colombia-U.S. Policy Rate Spread

	(1)	(2)	(3)	(4)	(5)	(6)
			$Loan_{f,b,yq}$			
i_{yq-1}*FX-$Funds_{b,yq-1}$*$Post_{yq}$		-81.450**	-73.523**	-77.064**	-83.296**	-126.016***
		(35.405)	(35.386)	(31.494)	(35.624)	(48.314)
i_{yq-1}*FX-$Funds_{b,yq-1}$		39.253	31.008	33.505	61.832**	68.681**
		(25.366)	(25.343)	(21.475)	(25.547)	(33.358)
FX-$Funds_{b,yq-1}$*$Post_{yq}$		-7.997	-12.628	-18.946	-20.728	-5.914
		(24.695)	(24.460)	(21.806)	(24.511)	(25.108)
FX-$Funds_{b,yq-1}$	-0.952***	-19.159*	-13.521	-12.200	-16.207	-19.789*
	(0.165)	(11.208)	(11.092)	(10.251)	(11.876)	(12.000)
Observations	895,247	895,247	895,247	895,247	895,247	895,247
R-squared	0.808	0.808	0.808	0.810	0.886	0.886
Firm*Bank FE	Yes	Yes	Yes	Yes	Yes	Yes
Macro Controls*Post	Yes	Yes	-	-	-	-
Bank Controls	Yes	-	-	-	-	-
Bank Controls*Post	No	Yes	Yes	Yes	Yes	-
FX-Funds*Macro Controls*Post	No	Yes	Yes	Yes	Yes	Yes
Year:Quarter FE	No	No	Yes	-	-	-
Industry*Year:Quarter FE	No	No	No	Yes	-	-
Firm*Year:Quarter FE	No	No	No	No	Yes	Yes
Bank Controls*i_{yq-1}*Post	No	No	No	No	No	Yes

This table shows how carry trade strategies by local banks impacts the reaction of bank credit to local monetary policy rate. The dependent variable, $Loan_{f,b,yq}$, is the log of total debt provided by bank b to firm f in year:quarter yq. i_{yq-1} is the lagged (by one quarter) local monetary policy rate. FX-$Funds_{b,yq-1}$ represents (lagged) bank FX-Funds (over Total Assets). Macro controls include the lagged values of annual GDP growth, of the CPI index and of the (log) Peso-US$ exchange rate. Bank Controls include lagged (by one quarter): ROA, log Total Assets, Common Equity (rescaled by Total Assets), Savings Deposits (rescaled by Total Assets) and Checking Deposits (over Total Assets). The sample includes only those companies that borrowed from at least two banks. Standard errors are double-clustered at the Bank*Industry and at the Firm level. *** p<0.01, ** p<0.05, * p<0.1.

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