International Financial Integration and Funding Risks: Bank-Level Evidence from Latin America*

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Abstract

Using a sample of over 700 banks in Latin America, we show that international financial liberalization lowers bank capital ratios and increases the shares of short-term funding. Following liberalization, large banks substitute interbank borrowing for equity and long-term funding, whereas small banks increase the proportions of retail funding in their liabilities, which have been particularly vulnerable to flight-to-quality during periods of financial distress in much of Latin America. We also find evidence that riskier bank funding in the aftermath of financial liberalizations is exacerbated by asymmetric information, which rises on geographical distance and the opacity of balance sheets.

Keywords: Bank Capital Structure, Financial Liberalization, International Capital Flows

JEL classification: F32, F36, G21

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1 Introduction

What are the implications of international financial liberalization for the funding structure of banks? While substantial macroeconomic research finds that international financial liberalization and bursts of foreign capital inflows increase the propensity for financial instability (Reinhart and Rogoff, 2008; Gourinchas and Obstfeld, 2012; Catão and Milesi-Ferretti, 2014), the microeconomic links between financial liberalization and financial instability are understudied. Existing papers on the implications of cross-border bank flows for financial stability focus on the asset side of bank balance sheets (Popov and Udell, 2012; Jordà et al., 2013; Taylor, 2015; Morais et al., 2015; Temesvary et al., 2015; Ongena et al., 2015; Correa et al., 2015; Bruno and Shin, 2015; Dinger and te Kaat, 2016; Hoffmann and Stewen, 2016), devoting scarce attention to the possibility that financial liberalization may affect bank stability through changes in banks' funding structures.

By changing the relative costs of the different types of funding, financial liberalization can lead to significant changes in the structure of bank liabilities. One mechanism is through enabling foreign investors to take positions in domestic banks. To the extent that information on the domestic bank is more costly or simply unavailable to the foreign investor, the degree of asymmetric information between the bank and its investor base rises. This lowers the cost of funding sources that are less sensitive to information asymmetries, such as debt, relative to those which are more sensitive to asymmetric information, such as equity (Myers and Majluf, 1984). Evidence from the behavior of broad stock price indices and bond spreads following major capital account liberalizations is consistent with this shift in the costs of debt relative to equity for emerging markets (Stulz, 1999; Bekaert and Harvey, 2000). It is therefore reasonable to hypothesize that capital account liberalization tends to raise the share of debt relative to equity in domestic banks' liabilities.[1]

In addition, financial liberalization can make banks more reliant on shorter-term debt. One mechanism is through the steepening of a country's yield curve around liberalization events. This can be due to lingering uncertainty on the longer-term sustainability of lib-

[1]An alternative theoretical channel is that financial liberalization intensifies the competition for input factors, such as labor (e.g., Cubillas and González, 2014). This may also lead to an overproportional reliance on debt insofar as the latter insures equity holders partly against failures of negotiations with suppliers of input factors, which, in turn, increases the bargaining position of shareholders (Sarig, 1998; Lee, 2011). A further theoretical mechanism is modeled by Innes (1990), who shows that high debt levels in the presence of asymmetric information reduce the incentives of private benefit taking.

eralization reforms (Calvo and Végh, 1999) and/or due to liberalizations that take place during periods of below-average global short-term interest rates. Another mechanism is highlighted in Flannery's (1986) work on the capital structure of firms: more limited capacity by new investors (in our case foreign investors) to distinguish between good and bad firms, makes good firms perceive their long-term debt as relatively underpriced, and thus lean towards issuing short-term debt.[2] Empirical support for a shift towards short-term debt due to such information frictions is provided in Johnson (2015) and Stohs and Mauer (1996). Against this body of theoretical and empirical evidence, it is thus reasonable to hypothesize that also the share of short-term to longer-term debt in banks' balance sheets is likely to rise in the wake of external financial liberalization.

This paper examines the evidence and above hypotheses regarding the effects of international financial integration on the patterns of equity and long-term funding (that is, debt with maturity above one year), using bank-level data from Latin America during 1995-2013. Latin America is particularly suitable for this investigation because extensive liberalization in external capital accounts was far-reaching and displayed considerable cross-country heterogeneity through the 1990s and 2000s, aiding identification of its effects on funding structures using panel data.[3] At the same time, restricting the sample to a single region like Latin America, helps filter out the effect of potentially powerful region-specific factors emphasized in Cerutti et al. (2015), which would call for more evolved and (arguably) less consensual model restrictions to help identification of regional factors. The main novelty of our analysis is to relate banks' stable funding ratios to international financial integration by applying dynamic panel data techniques and controlling for a wide variety of macroeconomic and bank-level factors. For this purpose, we construct a novel dataset matching bank-level data from Bureau van Dijk's Bankscope database for 17 Latin American countries and combine it with a very rich set of macroeconomic variables and the new index of financial openness by Fernández et al. (2015a), which has not previously been used in the attendant literature. Unlike previous indices of external financial openness or capital controls (such as that of Chinn and Ito, 2006), it distinguishes between

[2] See also Calomiris and Kahn (1991). Similarly, Rodrik and Velasco (1999) and Jeanne (2009) show that emerging market governments can signal better real returns by issuing short-term debt.

[3] In other emerging market regions, such as Asia, the process of capital account liberalization was less dynamic.

regulatory controls on capital inflows vs. capital outflows, and within those, provides a breakdown by type of flow. This distinction is clearly important as the purpose at hand is to evaluate the effects of regulatory changes in capital controls, actual external borrowing by domestic banks, and the composition of such borrowing.

In addition to being the first—to the best of our knowledge—to document the correlation between capital account liberalization and bank funding structures using bank-level data across Latin America, we also perform tests that examine whether these correlation patterns are consistent with the theoretical mechanisms reviewed above. First, we test the hypothesis that financial openness modifies the costs of the different types of funding by estimating the model separately for episodes of low and high US money market interest rates, as this interest rate gauges the potential reduction in costs associated with the substitution of foreign for domestic funding. Second, we explore the role of information asymmetry in shaping the dynamics of bank funding structures. We do so by allowing the decrease in stable funding ratios following international financial liberalization to be disproportionate when: (i) the bank has a riskier composition of assets, which we use as a proxy for the opaqueness of bank balance sheets and (ii) its foreign lenders are geographically more distant—an effect documented in distinct contexts by Hauswald and Marquez (2006), Mian (2006), and De Haas and Van Horen (2013). Third, we further explore the role of asymmetric information in the external liberalization-bank funding nexus by singling out the effects of liberalization on funding via retail deposits and interbank loans—funding types that can be withdrawn easily and, as a consequence, that are less sensitive to asymmetric information. In this specification, we also differentiate between large and small banks, since small banks are arguably more difficult to monitor by foreign investors and are usually less protected by implicit bail-out guarantees.[4] Fourth, as the value of bail-out guarantees is dependent on the institutional framework and the degree of financial stability, we also test the significance of institutional quality variables and of the incidence of crises for the behavior of the stable funding ratio and maturity structure following external liberalization.

Our results are as follows. First, we find financial liberalization to be associated with

[4]See Kang and Stulz (1997), who show that foreign investors, partly due to the presence of asymmetric information, are more likely to invest in large, well-known firms.

lower stable funding ratios, i.e., lower capital ratios and a stronger reliance on short-term funding. For instance, an increase in the liberalization index by one standard deviation reduces the average stable funding ratio by approximately 0.5-0.7 percentage points on impact and as much as 1.4-1.7 pp in the long-run. The economic significance of this results is highlighted by previous work which shows that even nominally small reductions in banks' stable funding ratios can increase the probability of bank distress disproportionately. For instance, the ECB (2015) underlines that a 1-pp increase in the Tier 1 capital ratios reduces the probability of distress relative to non-distress by 35-39% (see also Altunbas et al., 2014). Second, we show that this result is most pronounced during episodes of low US interest rates and of financial stability in the home country. Third, consistent with the presumption about the role of asymmetric information, we obtain overproportional effects in banks with more opaque balance sheets and for banks whose lenders are more distant. Fourth, we find liberalization to lead to higher interbank liabilities in large banks and to higher shares of retail deposits in small banks. As a result, large banks become more subject to rollover risks in the interbank markets, whereas smaller banks become more vulnerable to deposit withdrawal/bank run risks, which history has shown to be non-trivial in Latin America—and especially so in more institutionally fragile economies in the region.

These results speak to a scarce empirical literature that analyzes the determinants of banks' funding decisions (Song and Thakor, 2007; Berger and Bouwman, 2009; Dinger and von Hagen, 2009; Hahm et al., 2013; Craig and Dinger, 2014). Relative to these studies, we relate changes in funding structures to capital control regulations. This is important in the context of another strand of research that finds bank lending behavior, in particular during financial crises, to be contingent on the funding structure of banks (Dinger and te Kaat, 2016; Temesvary et al., 2015; Hoffmann and Stewen, 2016; Popov and Udell, 2012; Ongena et al., 2015). By finding that financial liberalization is associated with lower capital ratios and higher shares of more volatile short-term funding, this paper contributes to the macro literature that identifies international financial integration and foreign capital flows as important determinants of financial instability (Reinhart and Rogoff, 2008; Gourinchas and Obstfeld, 2012; Mendoza and Terrones, 2012; Catão and Milesi-Ferretti, 2014). The results presented here on the changes in the capital structure of banks in emerging markets can be seen as consistent with similar arguments regarding the more general capital structure shifts following capital account liberalizations in emerging economies. More specifically, they are consistent with results regarding corporate leverage ratios in emerging economies (Stiglitz, 2000; Booth et al., 2001; Lucey and Zhang, 2011), as well as with the results concerning the maturity structure of sovereign borrowing (Broner et al., 2013). Last but not least, our results add to the literature on the risk-taking channel of monetary policy in a cross-border environment (e.g., Babin, 2015; Ioannidou et al., 2015) by underlining that an expansionary monetary policy in the US disproportionately affects the funding structure of banks in peripheral economies and, thereby, their financial stability if the capital account is liberalized.

The remainder of this paper proceeds as follows. Section 2 describes our dataset. Section 3 lays out the empirical methodology and identification strategy. Section 4 presents the baseline results. In Section 5, we examine the role of asymmetric information in affecting banks' funding structures. Section 6 performs various robustness checks. Section 7 concludes.

2 Data

Our annual data spans the 1995-2013 period and 17 Latin American countries, namely Argentina, Bolivia, Brazil, Chile, Colombia, Costa Rica, Dominican Republic, Ecuador, El Salvador, Guatemala, Mexico, Nicaragua, Panama,[5] Paraguay, Peru, Uruguay and Venezuela.[6]

2.1 Bank-Level Data

Our bank-level variables are constructed from information provided in Bureau van Dijk's Bankscope database. We mostly include unconsolidated balance sheet data (i.e., Bankscope codes U1 and U2) because consolidated statements might be affected by foreign sub-

[5]Excluding Panama—which serves as a financial center—does not affect our estimates.

[6]Three Latin American countries (Cuba, Honduras and Puerto Rico) are not covered because of missing data on their degree of financial openness. We start our sample period in 1995 because both our bank-level data and the measure for capital account openness is not available before.

sidiaries.[7] After some data cleaning with regard to mergers and implausible observations (e.g., negative equity or liabilities), we obtain a sample of 8,982 bank-year observations, structured in an unbalanced panel of up to 766 banks that we track for up to 19 years.[8] Table 1 presents the number of financial institutions in every country of our sample over time. Bankscope coverage is lower for the 1990s relative to the 2000s, which results in a lower number of banks in our dataset for 1995-1999. As we will show in the sensitivity analysis presented in Section 6, we obtain qualitatively similar results for time sub-periods with a relatively constant number of banks.

We use this rich bank-level dataset to calculate our dependent variable as the share of equity and long-term liabilities (that is, debt liabilities with a maturity above one year) in total assets ($STABLEFUNDING$). The choice of this definition of the stable funding ratio is mainly driven by the theoretical arguments presented in the introduction, which predict that international financial integration lowers bank capital ratios and, additionally, leads to a shift in the maturity of bank liabilities towards short-term funding. Analyzing the effects of financial liberalization on $STABLEFUNDING$ ensures that we capture the theoretical predictions with regard to both equity and the maturity of debt in one variable.[9] We exclude retail deposits from this definition of stable funding because—as we will show in Section 5.3—they are less stable in Latin America than in developed economies, particularly during episodes of financial distress. All in all, $STABLEFUNDING$ captures in one variable the dynamics of equity ratios and banks' long-term funding attributable to financial liberalization. Consistent with the theoretical models, we expect this variable to decrease when a country liberalizes its capital account.

In our estimations, we also include variables that are likely to affect the funding structures of banks. These include bank size, defined as the logarithm of total assets ($SIZE$), the ratio of liquid over total assets ($LIQUIDITY$), the return on assets ($PROFITABILITY$) and the share of non-interest income over gross revenues ($NONINTERESTINCOME$).

[7] When banks only report consolidated statements, we include these in our regressions.

[8] We lose 686 observations because of the merger correction. Moreover, 451 implausible observations are dropped.

[9] This definition of stable funding is also consistent with recent Basel III regulations, which define equity and long-term debt as the nominator of the net stable funding ratio (NFSR).

Table 1: The Distribution of Banks in our Sample over Time

Country	1995	2000	2005	2010	2013
Argentina	62	99	80	79	56
Bolivia	6	12	14	15	12
Brazil	99	138	114	131	100
Chile	34	29	32	32	36
Colombia	28	33	29	62	72
Costa Rica	12	52	70	69	61
Dom. Republic	7	31	42	58	50
Ecuador	4	25	24	22	24
El Salvador	7	16	16	19	20
Guatemala	21	30	26	28	29
Mexico	27	46	53	90	132
Nicaragua	8	10	10	12	11
Panama	53	66	63	61	43
Paraguay	10	21	13	14	16
Peru	16	20	22	28	25
Uruguay	4	39	23	24	17
Venezuela	12	62	36	22	32
$\sum$	410	729	667	766	736

2.2 Macroeconomic Data

Our proxy for financial liberalization is the extent of restrictions on capital inflows, measured by the Schindler inflow index (Fernández et al., 2015a). It is a new de-jure index of financial liberalization, measuring the strength of capital controls imposed by national authorities based on the IMF's Annual Reports on Exchange Arrangements and Exchange Restrictions. The index is calculated from 1995 to 2013 as the average of ten disaggregated inflow restrictions on single asset categories and takes the values of zero (fully liberalized) to one. In our model, $LIBERALIZATION$ is calculated as (1-Schindler inflow index) because—due to this transformation—higher values represent financial liberalization, facilitating the interpretation of our main coefficients. There are three key advantages of this index. First, it is a de-jure measure of international financial integration. This is beneficial relative to de-facto measures (such as the amount of foreign assets or foreign liabilities to GDP) because the de-jure measure is predominantly determined at the aggregate policy level and is arguably more exogenous to the funding decisions of banks. Second, the index by Fernández et al. (2015a) reports the openness of capital in- and outflows separately. For the analysis of this paper, focusing on inflow restrictions is important because inflows of foreign capital are likely to be more important for the funding structure of banks compared with capital outflow restrictions.[10] Third, the possible disaggregation of the index in the various types of capital flows, such as money market flows, bond flows, financial credit flows, is beneficial compared with other liberalization measures because it allows us—in some specifications—to construct an individual liberalization index that mainly includes asset categories that benefit financial institutions.

In line with Dinger and von Hagen (2009) and Gropp and Heider (2010), the macroeconomic dataset of our paper also contains several additional variables that might affect the funding structures of banks. These include PPP adjusted per capita GDP (*PERCAPITAGDP*), the regulatory capital requirement (*CAPITALREQUIREMENT*),[11] the percent change in the consumer price index to control for the high inflation rates in many Latin American countries (*INFLATION*) and the real exchange rate (*REALEXCHANGERATE*).[12] Our expectation related to the sign of these controls is that banks' ratios of equity and long-term funding are lower the lower is the regulatory capital requirement, the higher is the inflation rate, the more overvalued is the real exchange rate and the lower is per capita GDP. Our macroeconomic dataset also includes the unemployment rate, stock market volatility, the rule of law, the regulatory reserve requirements and sovereign debt. Yet, as these variables turned out to be statistically insignificant in the regressions, we exclude them from the vector of macro controls in the regression specifications reported in the remainder of this paper. Table A.1 (Appendix) provides further specifics of the data.

2.3 Combined Dataset

As noted in the introduction, one innovation of our paper is to merge a bank-level dataset for Latin America with the new index on capital account openness constructed by Fernández et al. (2015a) and with a rich set of macroeconomic controls. Figure 1 provides an overview of the time-series evolution around major liberalization events, defined as a 0.25-reduction in either bond or financial credit or money market inflow restrictions, of the three most critical variables of this combined data—namely the average stable funding ratio (i.e., the sum of long-term liabilities and equity relative to total assets) of Latin

[10]However, in one of our sensitivity tests, we underline the robustness of our results by focusing on a liberalization index based on net capital flow restrictions.

[11]See Barth et al. (2001).

[12]See Darvas (2012a), Darvas (2012b), Darvas (2012c).

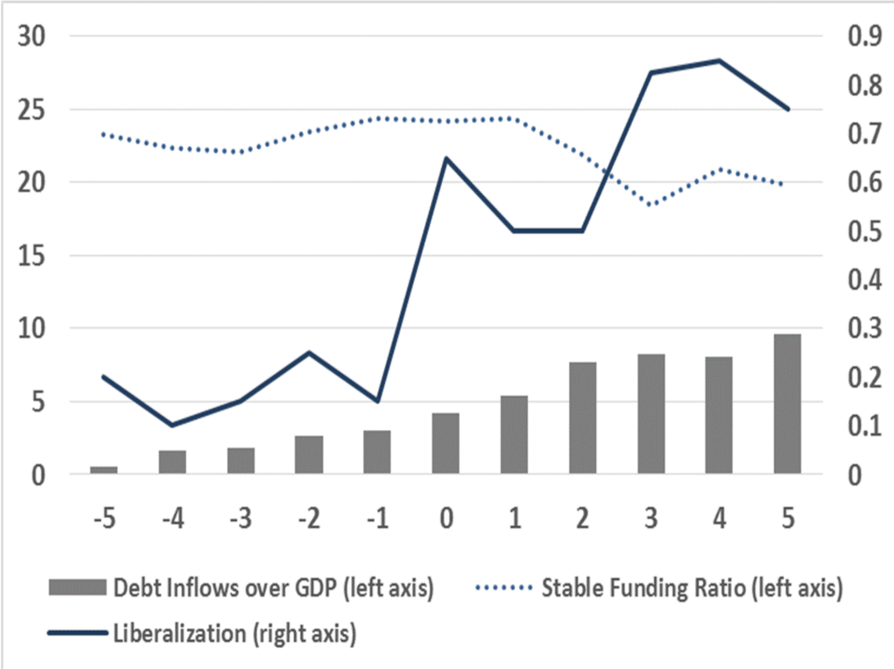

Figure 1: The evolution of the average stable funding ratio of banks (in % of total assets, dotted line), our continuous measure of financial openness (solid line) and the accumulated gross portfolio debt inflows (in % of GDP, bar chart) around major capital account liberalization events.

American banks, our continuous measure of financial openness and the accumulated gross portfolio debt inflows over GDP. The graph provides prima facie evidence that liberalization (i) leads to a sharp increase in debt inflows and (ii) in parallel, reduces banks' shares of stable funding.

After presenting the summary statistics of these and other variables of interest in the next section, the remainder of this paper will then dig into the prima-facie relationship between bank funding and financial liberalization plotted in Figure 1 using dynamic panel data models.

2.4 Summary Statistics

Table 2 summarizes the main descriptive statistics of the bank-level and macroeconomic variables in our model. The arithmetic mean of the stable funding ratio in our dataset is 27.15%, implying that the share of equity and long-term funding in total assets is more than a quarter. The variation of this variable, however, is very pronounced, with a 10th percentile of 8.77% and a 90th percentile of 62.69%.

We also control for various bank-level variables in our estimations. Table 2 shows that the average return on asset is equal to 1.81% and the ratio of liquid assets to total assets, calculated as (total assets-loans) / total assets, has a mean of about 47.2%. This result implies that the asset side of Latin American banks consists of a significantly lower share of loans compared with bank balance sheets in more developed countries.[13] Moreover, the average share of non-interest income over gross revenues in our sample is equal to 30.59%—a result that is similar to earlier research on advanced economies (e.g., DeYoung and Rice, 2004).

The values for per capita GDP in our sample also vary substantially with a 10th percentile of 5,770 USD and a 90th percentile of 16,520 USD. The average regulatory capital requirement is 9.71% and the average inflation rate equals 8.64%. Non-trivial inflation suggests that it is crucial to correct for changes in price levels, as they are likely to affect our estimates.

Table 2: Summary Statistics

	Obs.	Mean	SD	10th	Median	90th
STABLEFUNDING	8982	27.15	22.33	8.77	19.15	62.69
LIBERALIZATION	8981	0.66	0.32	0.15	0.80	1.00
SIZE	8982	6.21	2.13	3.50	6.17	8.97
PROFITABILITY	8950	1.81	4.86	-0.36	1.50	5.12
LIQUIDITY	8849	47.24	22.87	19.17	44.06	81.13
NONINTERESTINCOME	8849	30.59	42.74	1.29	25.82	70.48
PERCAPITAGDP	8981	10.79	4.11	5.77	10.53	16.52
CAPITALREQUIREMENT	8981	9.71	1.43	8.00	10.00	11.00
INFLATION	8895	8.64	11.30	1.48	5.66	16.21
REALEXCHANGERATE	8981	107.21	29.69	81.76	102.73	125.30

Note: The definitions, sources and units of these variables can be found in Table A.1 (Appendix).

[13] Compare Dinger and te Kaat (2016), who—for European banks—find a median loan-to-asset ratio of about 60%.

3 Empirical Methodology

3.1 The Exogeneity of Liberalization: A First Pass

Our argument of the causal effects of capital account liberalization on banks' funding structure relies on the exogeneity of capital control measures with respect to banks' funding decisions. Since banks' funding decisions may have a non-trivial effect on both actual external borrowing and policy makers' decision to liberalize the capital account, it is instructive to provide some institutional background on capital account liberalization in Latin America. Starting in the late 1980s, Latin American countries started reducing restrictions on external borrowing by financial and non-financial corporates.

Figure 2 displays the average degree of capital account openness in Latin America, Central and Eastern Europe and Asia during 1990-2013. The chart underlines that the liberalization process was most evolved in Latin America, especially compared with Asia, but also compared with Central and Eastern Europe.[14] While the pervasiveness of the controls and the extent of their reduction varied widely across the region, the general gradient was clear. In much of the region, it was motivated by a less pressing need to generate external trade surpluses to repay external debt in the wake of debt write-offs and debt settlement with foreign creditors, which started re-pulling capital back in from the early 1990s onwards. In countries with IMF programs, those were an additional prodding force. Another determinant was a global trend towards financial liberalization, which started in advanced countries—notably, the US and the UK—earlier in the 1980s. Furthermore, as argued by Brooks (2004), the political orientation of the incumbent government appears to have been a significant determinant of the decision for capital account liberalization. This encompasses the case of Mexico, where some domestic political consensus was finally forged by the newly formed technocratic government to advance with the country's membership into NAFTA. Since the freedom of capital movements was an important requirement of

[14] Several issues with Central and Eastern Europe motivate us to focus solely on the Latin American experience, rather than on both regions. First, capital account liberalization in Central and Eastern Europe was part of a more general process of institutional reforms, which implies that the region prior to the capital account liberalization had almost no tradition of market institutions, with most banks being state-owned. Second, the liberalization process was almost exclusively initiated immediately after most of the countries suffered a banking crisis in the mid to late 1990s, so that pre-liberalization stable funding ratios are affected by crises effects.

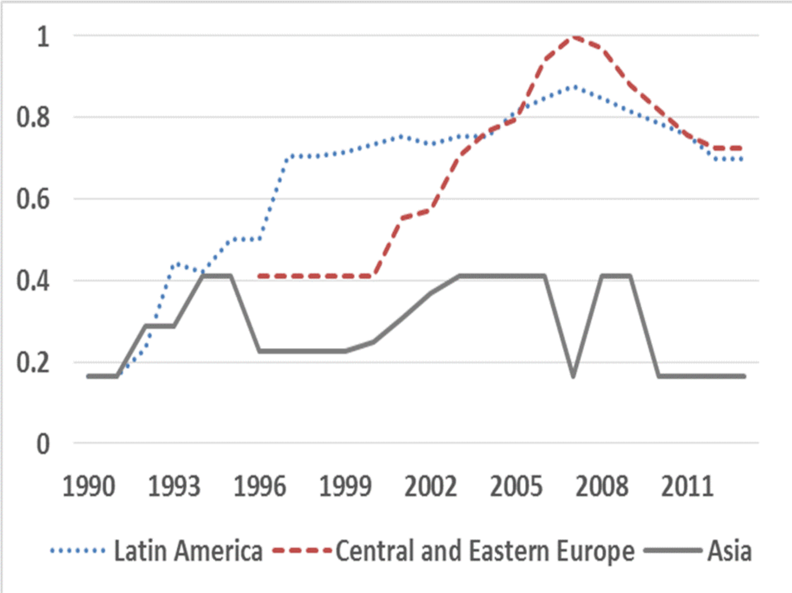

Figure 2: The degree of capital account openness in Latin America (dotted line), Central and Eastern Europe (dashed line) and Asia (solid line) during 1990-2013.

that trade treaty, the decision to join NAFTA was instrumental to the disbanding of the stringent system of capital controls, prevailing since the debt crisis in the 1980s. Elsewhere in the region, other idiosyncratic elements also played a role. This was the case in Brazil in the early 1990s when, under the liberal orientation of president Collor de Mello, trade and capital flows were liberalized as a political response to the inefficiency of domestic monopolies, aiming to grant nationals wider access to lower cost/higher quality imported goods as well as broader opportunities to allocate their savings.[15]

These considerations suggest that capital account restrictions are mainly exogenous to macroeconomic covariates, such as the domestic business cycle and the capital inflow cycle in different countries. However, policy makers in the emerging markets also frequently impose capital account restrictions for counter-cyclical reasons associated with overshootings and undershootings in external borrowing by residents, including the domestic banking sectors. If liberalization events indeed occur as a counter-cyclical response to the behavior of domestic agents, causal inference might be challenged. Yet, as recent work by Fernández et al. (2015b) based on the new capital control measures used in this paper underlines, capital controls have been strikingly **a**-cyclical in the broad cross-country panel that they examine.

[15]See Trubek et al. (2013).

Table 3: The Determinants of Liberalization

	(1) LIBERALIZATION	(2) DUMMY_LIBERALIZATION	(3) DUMMY_DELIBERALIZATION
PERCAPITAGDP (t-1)	-0.024	-0.037	0.060
	(-0.85)	(-1.41)	(1.01)
CAPITALREQUIREMENT (t-1)	0.018	0.056	0.009
	(0.60)	(1.30)	(0.15)
INFLATION (t-1)	-0.002	0.000	-0.001
	(-0.99)	(0.01)	(-0.14)
REALEXCHANGERATE (t-1)	0.001	-0.005	0.001
	(0.34)	(-1.61)	(0.13)
SPREAD (t-1)	-0.004	0.000	0.000
	(-1.50)	(0.41)	(0.13)
VIX (t-1)	-0.006	-0.005	0.007
	(-0.84)	(-0.54)	(0.70)
Year FE	Yes	Yes	Yes
Country FE	Yes	Yes	Yes
Obs	221	221	221
R-squared	0.705	0.505	0.533

This table presents regression results on the determinants of the degree of capital account openness. In particular, we regress (i) the degree of international financial integration, (ii) a dummy, being equal to one if a country liberalizes either its restrictions on bond or money market or financial credit flows by at least 0.25 and (iii) a dummy, being equal to one if a country restricts one of these inflows by at least 0.25 on the set of macro controls presented in Section 2 and the real interest rate spread vis-à-vis the US and the volatility index. Moreover, we add year and country dummies. The regressions are weighted by the number of banks in a country. The standard errors are clustered at the country-level and the t-statistics are shown in parentheses.

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

To evaluate whether these results also hold for our sample of countries, Table 3 shows the results of weighted cross-country regressions with the following dependent variables: (i) the overall degree of capital account openness ($LIBERALIZATION$), (ii) a dummy that measures major liberalization events and (iii) a dummy that measures de-liberalization events. The independent variables span well-known macroeconomic determinants that could potentially affect international financial liberalization, in addition to country and year fixed effects. The results are consistent with those of Fernández et al. (2015b) in that no macroeconomic variable affects the degree of international integration consistently (column (1)). Moreover, no domestic variable has an impact on major liberalization or de-liberalization events (columns (2) and (3)).

Consistent with the above evidence, we next turn to our baseline specification, wherein we treat changes in financial account liberalizations as broadly exogenous to banks' funding decisions. We further corroborate on the robustness of this assumption and buttress our inferences by instrumenting changes in capital account openness with a government's partisanship indicator and a dummy variable to indicate when a country is under an IMF-sponsored adjustment program, in which case liberalization decisions are more plainly dictated by external conditionality.

3.2 Econometric Specification

We explore the relationship between capital account liberalization and stable funding ratios using the following model:

$$\begin{aligned} STABLEFUNDING_{i,t} &= \alpha_i + \gamma * STABLEFUNDING_{i,t-1} + \beta * LIBERALIZATION_{j,t} \\ &\quad + \theta * X_{i,j,t} + \varepsilon_{i,t} \end{aligned} \tag{1}$$

Our dependent variable is the ratio of equity and long-term liabilities (that is, debt liabilities with a maturity above one year) relative to total assets of bank i at time t ($STABLEFUNDING$).[16] As is apparent from Figure 1, the aggregate behavior of this variable is suggestive of non-trivial autocorrelation. It therefore seems important to allow for a lagged dependent variable on the right hand side of equation (1) to help capture the time-series dynamics of banks' funding structures.[17] The coefficient α_i is an individual bank intercept and the vector X includes the bank-level and macroeconomic controls, listed in Table 2, as well as—in some specifications—their interactions with categorical variables postulated by theory, as discussed below. The main coefficient of interest in the following analysis is β, which measures the short-run impact of financial liberalization on the stable funding ratio of banks. The long-run effect is given by $\frac{\beta}{1-\gamma}$.

Estimating equation (1) with OLS yields inconsistent estimates in the presence of individual bank-specific effects. If we simply replace pooled OLS with fixed effects regressions, the estimates may also be non-trivially biased by the presence of the lagged dependent variable once the panel's time series dimension is not too large (Nickel, 1981). To overcome these drawbacks, we make use of three distinct estimation procedures that all yield consistent coefficients: First, we estimate the equation with the estimator proposed by

[16] Theory predicts financial liberalization to lead to (i) a decrease in capital ratios and (ii) to shorter-term liabilities. Defining the dependent variable as we do ensures that we capture both of the theoretical predictions in one variable.

[17] This is standard, among others, in Faulkender et al. (2012).

Anderson and Hsiao (1981). It estimates equation (1) in first differences and instruments the lagged dependent variable with its lagged levels.[18] However, as shown by Arellano and Bover (1995), the Anderson-Hsiao estimator magnifies gaps in unbalanced panels and often estimates coefficients imprecisely. Therefore, we also estimate (1) with the Blundell-Bond system GMM estimator (Blundell and Bond, 1998)[19], which uses both the variable levels as instruments for the equation in first differences and, additionally, first differences of the variables as instruments for the variables in levels. The existing literature on the determinants of firms' liability structures shows that the Blundell-Bond estimator is superior to the Arellano-Bond estimator (Arellano and Bond, 1991), in particular, because of the high persistence of the dependent variables (e.g., Faulkender et al., 2012; Flannery and Hankins, 2013). Finally, we also address any remaining endogeneity concerns about $LIBERALIZATION$ by estimating the Blundell-Bond estimator, but instrumenting the degree of capital account openness with a government's partisanship indicator and an IMF program dummy, both of which are broadly exogenous to financial liberalization and likely to be significant drivers of the latter (see the discussion in Section 3.1).

For the Blundell-Bond regressions, we instrument the regressors with four lags of their levels and first differences, respectively. Restricting the instruments to four lags is important because the number of instruments for this estimator increases quadratically in T and, therefore, can become very large, overfitting endogenous variables (Roodman, 2009b). In our specifications, we also correct the standard errors by the procedure proposed by Windmeijer (2005). This procedure addresses the potential downward bias of the standard errors that arises when using a large number of instruments in a regression. Its application makes our t-statistics more conservative, leading to more reliable inference.

[18] We restrict the number of lags to two. However, we obtain qualitatively similar results for different lag lengths.

[19] We rely on the xtabond2 command in Stata (Roodman, 2009a) to estimate these regressions.

4 Results

4.1 Baseline Results

In Table 4, we present the results for the baseline regression specification of equation (1). We start—in column (1)—with the results for the Anderson-Hsiao estimator. In column (2), we also add macroeconomic controls to the set of bank-level controls. In columns (3) and (4), we present the results for the Blundell-Bond estimator and in columns (5) and (6), we additionally instrument $LIBERALIZATION$ with two exogenous instruments—a government's partisanship indicator and an IMF program dummy.

The results of all these specifications underline that capital account openness is associated with lower stable funding ratios. In particular, an increase in the liberalization index by one standard deviation (about 0.32 in our sample) reduces the share of equity and long-term funding relative to total assets on impact by 0.5-0.7 pp. The long-run effect ranges between 1.4 and 1.7 pp, as can be gauged by dividing the coefficient of $LIBERALIZATION$ by (1 - autoregressive coefficient). These are economically significant effects since other work shows that even smaller reductions in banks' stable funding ratios can greatly increase the probability of bank distress. More specifically, in a logit regression framework, the ECB (2015) finds that a 1-pp increase in the Tier 1 capital ratios reduces the odds ratio (that is, the probability of distress relative to non-distress) by 35-39% (see also Altunbas et al., 2014).

The only bank-level variables that affect the funding structures significantly are the profitability and the size of banks. Less profitable and larger banks have lower stable funding ratios, a result that is consistent with earlier findings (e.g., George, 2015; Gropp and Heider, 2010). Moreover, in line with our expectations, we find that higher per capita income, lower inflation and a real devaluation are associated with higher equity ratios and more long-term funding in some specifications. Therefore, in contrast to Gropp and Heider (2010), who find macroeconomic factors to be insignificantly associated with changes in bank funding structures in the US and Europe, we identify an important role of macroeconomic variables in affecting the dynamics of bank funding in Latin America.

Table 4: Results for the Baseline Model

	Anderson-Hsiao		Blundell-Bond		Blundell-Bond / IV	
	(1) STABLEFUNDING	(2) STABLEFUNDING	(3) STABLEFUNDING	(4) STABLEFUNDING	(5) STABLEFUNDING	(6) STABLEFUNDING
STABLEFUNDING (t-1)	0.421	0.420	0.661***	0.679***	0.653***	0.669***
	(1.46)	(1.46)	(14.45)	(16.86)	(13.59)	(16.18)
LIBERALIZATION	-2.141**	-2.103**	-1.408*	-1.619**	-1.840*	-1.729**
	(-2.56)	(-2.44)	(-1.67)	(-2.49)	(-1.78)	(-2.51)
SIZE	-6.197***	-6.609***	0.267	-0.162	-0.026	-0.202
	(-5.25)	(-4.72)	(0.98)	(-0.73)	(-0.09)	(-0.85)
PROFITABILITY	0.160	0.161	0.114	0.106	0.183*	0.119
	(1.48)	(1.45)	(1.14)	(1.24)	(1.74)	(1.36)
LIQUIDITY	-0.045	-0.039	0.014	-0.000	0.011	0.004
	(-0.62)	(-0.54)	(0.63)	(-0.02)	(0.46)	(0.17)
NONINTERESTINCOME	-0.000	0.000	-0.005	-0.001	0.000	-0.003
	(-0.00)	(0.00)	(-0.56)	(-0.12)	(0.04)	(-0.33)
PERCAPITAGDP		0.314		0.129**		0.148**
		(0.84)		(2.17)		(2.41)
CAPITALREQUIREMENT		-0.267		0.262		0.160
		(-0.57)		(1.13)		(0.64)
INFLATION		0.053		-0.059***		-0.054***
		(1.38)		(-3.06)		(-2.63)
REALEXCHANGERATE		0.023		-0.012*		-0.013*
		(1.03)		(-1.78)		(-1.80)
Obs	2865	2865	6660	6646	6660	6646
p (Hansen statistic)	0.32	0.31	0.21	0.10	0.50	0.06

In our baseline specifications, we regress the stable funding ratio, i. e., the share of equity and long-term funding relative to total assets on its own lag and a measure for financial liberalization that takes the values from 0 (capital account closed off) to 1 (fully liberalized). In the regressions, we also include a vector of bank - specific controls and, in some specifications, macroeconomic control variables. In column (1)-(2), we estimate the dynamic model with the Anderson-Hsiao estimator (Anderson and Hsiao,1981). Columns (3)-(4) make use of the Blundell-Bond estimator (Blundell and Bond, 1998), using four lags of the variables as instruments. In columns (5) and (6), we instrument LIBERALIZATION not by its own lags, but an IMF program indicator and a government's partisanship dummy. In columns (3)-(6), we apply Stata's xtabond2 command (Roodman, 2009a). The t-statistics that make use of Windmeijer (2005) corrected standard errors are shown in parentheses and p (Hansen statistic) provides the p values for the Hansen test of overidentification restrictions.

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

Except for the Anderson-Hsiao estimator, the lag of the stable funding is highly statistically significant with a coefficient between 0.6 and 0.7. This implies that the adjustment speed of the stable funding ratio (1-autoregressive coefficient) in our sample ranges between 0.3 and 0.4. An adjustment speed of this magnitude is consistent with that obtained by Faulkender et al. (2012) for target leverage ratios and suggests that stable funding ratios adjust quickly.

The above results indicate that the economic significance of $LIBERALIZATION$ is broadly robust to the chosen econometric estimator. Further, columns (5) and (6) show that instrumenting $LIBERALIZATION$ with a government's partisanship indicator and an IMF program dummy does not yield results that are significantly different from those relying on the exogeneity of capital controls, consistent with the discussion of section 3.1. As for which estimator is to be preferred, it is worth remarking that the Blundell-Bond estimator has been shown to be superior to the Anderson-Hsiao estimator, especially in unbalanced panel datasets (e.g., Blundell and Bond, 1998).

As previous research finds the Blundell-Bond estimator to be first choice in capital structure regressions (e.g., Faulkender et al., 2012), we will stick to that choice in all subsequent specifications. But again, the key finding that international financial integration reduces the share of equity and long-term funding relative to total assets and, therefore, increases banks' reliance on short-term funding is robust to that choice.

4.2 Controlling for US Monetary Policy

In this sub-section, we further expand the baseline analysis by testing the hypothesis that financial openness affects the funding structures of banks differently during periods of high vs. low money market interest rates in the world's main financial center—the US. Variations in US money market interest rates gauge the potential reduction in banks' funding costs associated with the substitution of foreign for domestic funding. Thus, the effects of financial liberalization might be expected to be stronger the lower the US interest rate. Since the US interest rate is mainly driven by the stance of monetary policy in the United States, it is broadly exogenous to Latin America and allows us to analyze the interaction of capital account liberalization and monetary policy in the US. Moreover, we are able to address the interaction between global financial cycles and domestic financial stability through the lens of capital account openness. Yet, an important econometric consideration in measuring these effects is that US money market rates have hovered around at the zero lower bound for several years and the effects of US monetary policy on banks' funding decisions in emerging markets may be rather more dependent on quantitative easing aspects of US monetary policy. To capture the latter as well, we use the shadow interest rate calculated by Wu and Xia (2016), which allows interest rates to turn negative when the zero lower bound hits and quantitative effects of US monetary policy may predominate. Using this indicator, we define a low US interest rate episode as one in which this shadow rate is below the median in our sample. When analyzing the impact of global interest rates, we account for the fact that the access of foreign banks to global low-interest funding is also likely to depend on the stability of the local financial system. During crises, foreign investors are likely to restrict their lending to banks in emerging economies and, as a result, stable funding ratios might be affected less during these episodes. Thus, the interaction between financial liberalization and monetary policy in the US is finally also tested by excluding observations coinciding with either a local or a global financial crisis, as defined by Laeven and Valencia (2013).

In columns (1)-(2) of Table 5, we find evidence that the impact of financial liberalization on bank funding structures is more precisely estimated (with a statistical significance at the 5% and 10% level) when US interest rates are low. Moreover, the higher estimated coefficient (0.72 vs. 0.54) on the lagged dependent variable points to a higher long-run effect during episodes of low US rates: The long-run effect of a 1-standard deviation increase in financial integration is 1.6 during low US rate episodes (column (1)) and significantly smaller and only equal to 1.0 during high interest rate episodes (column (2)). It might also be conjectured that the strength of this effect is contingent on the degree of financial stability in the host country. During financial crises, global investors are likely to restrict their lending to banks in Latin America even if US interest rates are low; hence, stable funding ratios might be affected less during such episodes. Indeed, the results reported in columns (3)-(6) show that during financial crises and low US money market rates, external financial liberalization does not affect stable funding ratios significantly: the relevant coefficient remains assuredly negative, but is not statistically significant at conventional levels.

Table 5: Controlling for US Monetary Policy

			Financial Crisis		*no Financial Crisis*	
	low US rate	**high US rate**	**low US rate**	**high US rate**	**low US rate**	**high US rate**
	(1) STABLEFUNDING	(2) STABLEFUNDING	(3) STABLEFUNDING	(4) STABLEFUNDING	(5) STABLEFUNDING	(6) STABLEFUNDING
STABLEFUNDING (t-1)	0.716^{***}	0.543^{***}	0.741^{***}	0.468^{***}	0.478^{***}	0.616^{***}
	(13.56)	(9.56)	(15.56)	(5.67)	(3.17)	(10.42)
LIBERALIZATION	-1.427^{*}	-1.487	-0.450	-1.168	-2.693^{**}	-1.131
	(-1.70)	(-1.40)	(-0.43)	(-0.63)	(-2.02)	(-0.90)
Bank-Level Controls	YES	YES	YES	YES	YES	YES
Macro Controls	YES	YES	YES	YES	YES	YES
Obs	3737	2909	2920	859	817	2050

In all these regressions, we estimate the baseline model separately for episodes of low and high US money market interest rates, as these gauge the reduction in costs associated with the substitution of foreign for domestic funding. As even during low US rate episodes, the access to foreign funding depends on whether or not there is a financial crisis, we additionally, in columns (3) - (6), exclude and include, respectively, years of financial uncertainty. In all columns, we regress the stable funding ratio,i.e., the fraction of equity and long-term funding relative to total assets on its own lag and a proxy for financial liberalization that takes the values between 0 (capital account closed off) and 1 (fully liberalized). We also add a set of bank- specific and macro controls. The regressions are implemented with Stata's xtabond2 command (Roodman, 2009a), using 4 lags of the variables as instruments. The t-statistics make use of Windmeijer corrected standard errors and are presented in parentheses.

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

In contrast, the combination of low US interest rates, a more open capital account and sound financial conditions reduces banks' shares of equity and long-term funding. The related coefficient of *LIBERALIZATION* is equal to -2.693 and significant at the 5% level. In a nutshell, we show that the strength of the effects of external financial liberalization on banks' funding structures are influenced by the stance of monetary policy abroad and financial stability at the host country. When both money market rates in international financial centers and financial instability are low, banks in liberalized economies reduce the shares of stable funding. In contrast, international financial liberalization has more muted effects on the banks' funding structures during higher US money market rates and/or financial distress. In other words, the funding structure of Latin American banks is affected most when local banks have access to relatively inexpensive international sources of funding. This result is consistent with our conjecture that liberalization affects the funding structure of banks by changing the relative costs of the different types of funding. It adds to the literature that explores the risk-taking channel of monetary policy in a cross-border environment (e.g., Babin, 2015; Ioannidou et al., 2015) by underlining that an expansionary monetary policy in the US disproportionately affects the funding structure of banks in peripheral economies—and, thereby, their financial stability—if the capital account is liberalized.

5 The Role of Asymmetric Information

Our empirical tests are predicated on the assumption that changes in capital controls affect banks' access to low-interest funding abroad and, thus, change the costs of the various types of bank funding. However, as noted in the introductory section, the *relative* costs of alternative types of funding following liberalization are likely to be a function of the degree of asymmetric information between lenders and borrowers—with stable funding (equity and long-term liabilities) staying relatively expensive and short-term funding that is less sensitive to information asymmetries becoming cheaper. In this section, we explore whether our dataset is consistent with the role of asymmetric information in influencing the post-liberalization dynamics of banks' funding decisions. In particular, we expect that—since short-term debt is least affected by asymmetric information, followed by longer-term debt and then by equity—the costs of equity and long-term debt relative to short-term debt costs are increasing in asymmetric information. Hence, higher information asymmetries between banks and their investors should lower bank capital ratios and increase their short-term funding. Our first test evaluates whether the decrease in stable funding ratios is disproportionate in banks with opaque balance sheets and lenders that are more distant. Our second test gauges the effects of financial liberalization on the ratio of retail deposits and interbank loans—funding types that can easily be withdrawn and, as a consequence, are less influenced by asymmetric information. In this connection, we differentiate between large and small banks because in general small banks are subject to higher information asymmetries—in part because they are also less protected by implicit bail-out guarantees.[20] Since the value of these implicit guarantees and, therefore, the degree of information asymmetries is dependent on the institutional quality and whether or not there is a financial crisis, our third test explores the effect of crises and the quality of institutions for stable funding ratios, retail deposits and interbank loans.

5.1 Opaqueness and Distance

To explore the direct role of asymmetric information in affecting the impact of financial liberalization on banks' funding structures, we present two tests. First, we examine whether the effect of liberalization is most pronounced in banks with opaque balance sheets. Opaqueness generally aggravates the monitoring abilities of lenders and increases the information asymmetries between banks and their investors. Second, we explore the impact of distance between banks and their investors in influencing the changes in banks' funding modes. Following Hauswald and Marquez (2006), Mian (2006) and De Haas and Van Horen (2013), we assume that the geographical distance inhibits the monitoring capacities of foreign lenders and, hence, increases information asymmetries.

We start by enabling the liberalization index to interact with the ratio of impaired loans relative to equity. While this ratio is a frequently used measure for the credit risk-taking incentives of banks, we follow Jungherr (2016) in presuming that a higher share of impaired loans generally signals that the bank is prone to funding more opaque projects,

[20]For a similar argument, see Kang and Stulz (1997), who show that foreign investors, partly due to the presence of asymmetric information, are more likely to invest in large, well-known firms.

whose values are subject to substantial degrees of asymmetric information (and, hence, whose recovery of principal and interest, once they fall in default, is also subject to greater uncertainty). If so, we should observe an additional effect of $LIBERALIZATION$ on banks with high impaired loans ratios.

The first two columns of Table 6 support this hypothesis: whereas the short-run effect of an increase in our measure of financial integration by one standard deviation on banks at the 25th percentile of the distribution of impaired loans to equity is equal to -0.57 pp, its effect on banks with higher asymmetries at the 75th percentile is equal to -0.75 pp (column (2)). This result suggests that the interaction term is not only statistically significant, but also economically important.

Table 6: Opaqueness and Distance

	(1) STABLEFUNDING	(2) STABLEFUNDING	(3) STABLEFUNDING	(4) STABLEFUNDING
STABLEFUNDING (t-1)	0.654***	0.653***	0.552***	0.630***
	(13.25)	(14.75)	(8.52)	(11.72)
LIBERALIZATION	-2.157**	-2.027***	20.308***	7.004
	(-2.35)	(-2.67)	(2.82)	(1.30)
LIBERALIZATION × RISK	-0.045*	-0.041*		
	(-1.83)	(-1.67)		
LIBERALIZATION × DISTANCE			-2.034***	-0.722
			(-2.87)	(-1.37)
RISK	0.031	0.028		
	(1.57)	(1.49)		
DISTANCE			1.976***	0.908*
			(3.25)	(1.85)
Bank-Level Controls	YES	YES	YES	YES
Macroeconomic Controls	NO	YES	NO	YES
Obs	5111	5111	2863	2849

In these regressions, we test whether financial liberalization affects the funding structure of banks through changes in asymmetric information. The dependent variable is the stable funding ratio, i.e., the share of equity and long-term funding relative to total assets that we regress on its own lag and a measure for financial liberalization that takes values from 0 (capital account closed off) to 1 (liberalized). Moreover, we interact the liberalization measure with the riskiness of banks (columns 1, 2) and the average weighted distance of the banking sectors vis-à-vis their international counterparties (columns 3, 4). We also include bank-level and macroeconomic controls in some specifications. All of the Blundell-Bond regressions are implemented with Stata's xtabond2 command (Roodman, 2009a), using 4 lags of the variables as instruments.The t-statistics make use of Windmeijer(2005) corrected standard errors and are shown in parentheses.

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

To illustrate the role of informational asymmetries, we perform an additional test, which explores whether the effect of financial liberalization is also most pronounced in countries in which foreign lenders are on average more distant. This test is motivated by the assumption that distance inhibits the monitoring abilities of foreign lenders and, therefore, increases informational asymmetries. For this analysis, we rely on the consolidated banking statistics of the Bank for International Settlements, which publishes the outstanding amounts of banking sector liabilities vis-à-vis the banking sectors in foreign countries. We use this data to calculate a weighted measure of the distance between the local banking system and its international lenders as follows:

$$DISTANCE = \frac{\sum(distance_{jk} * outstandingliabilities_{jkt})}{\sum outstandingliabilities_{jkt}}, \tag{2}$$

where *outstandingliabilities* is the amount of interbank liabilities of country j vis-à-vis country k at time t and *distance* is the distance between these two countries j and k.
The variable *DISTANCE* is strongly centered around 7.5 because most banking sectors in Latin America maintain close relationships to the US, whose distance from most countries in Latin America is about 7,500 kilometers.[21] However, several countries (i.e., Argentina, Chile, Panama and Uruguay) have an average distance of more than 10,000 kilometers, whereas numerous countries (i.e., Costa Rica, Dominican Republic, El Salvador, Guatemala and Nicaragua) are also relatively close to their international lenders. In the following test, we explore whether these two country groups differ in their sensitivity with respect to financial liberalization. By interacting *DISTANCE* with the liberalization index, we are able to test whether liberalization differently affects those countries that are geographically close to their international lenders relative to the set of geographically distant countries.[22] Column (3) of Table 6 provides further evidence for our hypothesis that financial liberalization affects the liability side of banks mainly through higher informational asymmetries. In this column, we obtain an amplified effect of international financial integration in banking sectors with lenders that are on average more distant. This

[21] Although Mexico is very close to the US, Mexican banks to large extents maintain funding relationships with Spanish banks, increasing the average distance to its lenders to about 7,500 kilometers.

[22] The number of observations in this specification declines because we only explore the effects for these two sub-groups.

means economically that the stable funding ratios in banking sectors with an average distance of 6,000 kilometers (e.g., Costa Rica) even increase by 2.59 pp on impact when the liberalization index improves by one standard deviation.[23] In contrast, the stable funding ratios in banking sectors with an average distance of 12,000 kilometers (e.g., Panama) decrease in a highly significant manner by 1.31 pp. In column (4), we also add the vector of macroeconomic controls to the distance regressions. Once, we introduce these covariates, the interaction term of $LIBERALIZATION$ and $DISTANCE$ turns insignificant. Yet, there is a clear reason for that: as countries in Southern Latin America (e.g., Argentina and Chile)—that are more distant from the pool of their potential borrowers—are on average richer, the correlation between $DISTANCE$ and $PERCAPITAGDP$ is equal to 60%, leading to multicollinearity issues. Excluding per capita GDP from the set of macro controls again yields a negatively significant interaction term at the 5% level.[24]

To sum up, the results presented in this section provide empirical support that increased informational asymmetries are an important channel for the effect of international financial liberalization on banks' funding structures. This evidence is consistent with the classical pecking order arguments (Myers and Majluf, 1984).

5.2 Bank Size and the Costs of Asymmetric Information for Different Types of Funding

We now further explore the role of informational asymmetries by focusing on the cross-country and cross-bank differences in the informational sensitivity of various liabilities. More specifically, we explore the variation in the informational sensitivity of retail deposits and interbank funding—funding types that can quickly be withdrawn and, therefore, are less affected by asymmetric information. Especially retail deposits in large banks and in countries with well-developed institutions have been shown to be less sensitive to information because of the existence of explicit and implicit deposit and government guarantees.[25] On the contrary, retail deposits in small banks and in banking systems character-

[23]The standard deviation of $LIBERALIZATION$ is close to 0.32. The economic effect is obtained by making use of the $LIBERALIZATION$ and $LIBERALIZATION * DISTANCE$ coefficients as follows: 2.59=0.32*(20.308-2.034*6).

[24]This result is not reported to keep the table crispier, but is readily available from the authors upon request.

[25]See Boyd and Runkle (1993) for an overview.

ized by low levels of institutional quality are typically prone to informational asymmetry issues and, thus, less stable. In this subsection, we start by focusing on the cross-bank variation in terms of size. Our goal is to examine whether liberalization differently affects the funding structure of small and large banks, reflecting the differences in information sensitivity.

Following Cetorelli and Goldberg (2012), we introduce a dummy ($DUMMY-LARGE$) that is equal to one if a bank's amount of total assets is above the 95th percentile of the annual distribution and zero if total assets are below the 90th percentile. Then, we enable $LIBERALIZATION$ to interact with $DUMMY-LARGE$ in all of the following regressions. Additionally, we replace the stable funding ratio with the proportion of customer deposits (column (3) and (4)) and interbank liabilities (column (5) and (6)) relative to total assets in some specifications.

In column (1) and (2) of Table 7, none of the interaction terms is significant. This means that liberalization does not have heterogeneous effects on the stable funding ratios of small and large banks, respectively, suggesting that the different degree of information asymmetries between large and small banks does not affect our baseline results.

However, we obtain heterogeneous implications for large and small banks in terms of the composition of non-stable funding measured by the proportion of retail and interbank deposits. For the customer deposits, our estimates deliver a negative and significant interaction term. This implies that small banks—in contrast to large banks—increase their shares of retail deposits. In contrast, large institutions fund larger proportion of their balance sheets with interbank loans. This is an implication of the positively significant interaction term in columns (5) and (6).

Economically, a 0.32-increase (about one standard deviation) in the liberalization index leads large banks to raise the share of interbank funding by 0.4-0.7 pp in the short-run, whereas they reduce their reliance on customer deposits by approximately 0.5 pp (column (4)). In contrast, small banks increase their shares of retail deposits on impact by 1.4 - 1.5 pp and they reduce the share of interbank funding by about 0.5 pp.

Table 7: Bank Size and the Various Types of Funding

	(1) STABLEFUNDING	(2) STABLEFUNDING	(3) RETAIL	(4) RETAIL	(5) INTERBANK	(6) INTERBANK
STABLEFUNDING (t-1)	0.626***	0.654***				
	(13.53)	(16.07)				
RETAIL (t-1)			0.718***	0.735***		
			(19.51)	(23.45)		
INTERBANK (t-1)					0.651***	0.644***
					(19.35)	(21.90)
LIBERALIZATION	-1.982*	-1.923**	4.380***	4.795***	-0.867	-1.541**
	(-1.87)	(-2.34)	(3.22)	(4.74)	(-1.06)	(-2.44)
DUMMY_LARGE × LIBERALIZATION	1.190	0.969	-3.818**	-5.644***	2.059*	3.630***
	(0.65)	(0.56)	(-2.02)	(-2.94)	(1.85)	(3.34)
DUMMY_LARGE	0.374	2.069	5.465**	4.057*	0.819	-1.833
	(0.14)	(0.88)	(2.22)	(1.72)	(0.64)	(-1.45)
Bank-Level Controls	YES	YES	YES	YES	YES	YES
Macroeconomic Controls	NO	YES	NO	YES	NO	YES
Obs	6190	6177	8389	8292	6629	6531

In these regressions, we determine the effect of financial liberalization not only on the stable funding ratios, but also on the shares of retail and interbank funding. Moreover, we interact the financial liberalization measure that takes the values between 0 (capital account closed off) and 1 (fully liberalized) with a dummy that is one for banks in the top 5 % of the distribution of bank size and zero if a bank is in the lowest 90 %. This extension is important because we expect the effects on retail deposits and on interbank borrowing to be dependent on bank size. We also include the lagged dependent variable and bank-level and macro controls in some specifications. All of the Blundell-Bond tests are implemented with Stata's xtabond2 command (Roodman, 2009a), using 4 lags of the variables as instruments.The t-statistics use Windmeijer (2005) corrected standard errors and are shown in parentheses.

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

In sum, the results illustrate the role of the variation in the informational sensitivity of retail deposits and interbank funding, presenting further evidence for the importance of asymmetric information as a channel underlying the effect of international financial liberalization. These results further corroborate the existence of significant substitution effects between different liabilities when international financial liberalization takes place. For large banks that are subject to lower degrees of asymmetric information, capital account openness promotes interbank lending from abroad. This is likely to induce them to reduce the interest rates on retail deposits, thereby, leading to deposit flows to smaller banks. Consequently, whereas large banks benefit directly from financial integration through a better access to wholesale funding from abroad, small banks benefit indirectly because they have an increased share of retail deposits.

Increased interbank funding is generally associated with higher bank funding risks because interbank loans can easily be withdrawn and they are particularly volatile during financial crises (e.g., Bednarek et al., 2015). However, it might be that retail deposits—especially in countries without sound deposit insurance schemes and weaker institutions—are also volatile during crises and, hence, increase the funding risks of banks relative to equity and long-term funding. We discuss this possibility next.

5.3 The Role of Bail-out Guarantees During Crises

Having shown that bank size is a significant cross-bank driver of the informational sensitivity of bank funding in our sample, this subsection focuses on the cross-country dimension in this variation by exploring the value of implicit and explicit government guarantees. The stability of the various types of funding is likely to be contingent on the existence of good institutions and reliable deposit insurance schemes because they constitute an implicit bail-out guarantee, minimizing the implications that are associated with asymmetric information. As the value of these guarantees is mostly obvious in times of financial turmoil, we focus on the effects of financial liberalization on the funding structure of banks during banking sector crises.

In the following tests, we interact our measure of international financial integration with a crisis dummy that is equal to one if there is either a global or a local banking crisis (Laeven and Valencia, 2013) in the respective country-year pair. Moreover, we estimate the regressions separately for countries without a deposit insurance scheme, countries with a deposit insurance and good institutions and countries with a deposit insurance scheme but weak (below the 25th percentile of the annual distribution of the rule of law) institutions, respectively.[26]

Columns (1)-(3) of Table 8 stress that in countries without a deposit insurance, financial liberalization during "good" times increases the likelihood of banks' access to interbank funding at a 1% level (t-ratio of 3.03). In contrast, the share of retail deposits is not affected, as can be seen from the very low t-ratio of -0.4 on the same variable.[27] During crises, banks in countries without deposit insurance have lower interbank liabilities; however, financial crises do not affect the access to retail customer deposits. In countries with a deposit insurance and a good institutional framework (columns (4)-(6)), financial crises do not affect stable funding ratios, retail deposits or the shares of wholesale funding, indicated by insignificant interaction terms. These results also underline that retail funding is a stable form of funding in Latin American countries with a strong institutional framework. This result, however, does not hold for countries with weaker institutions (columns (7)-(9)). In these countries, depositors withdraw most of their savings during financial crises. In particular, whereas an increase in the liberalization index by one standard deviation increases the share of retail funding in the short-run by 2.37 pp at sound economic times, its effect during crises shrinks to only 0.69 pp,[28] indicated by a negatively significant interaction term. The stable funding ratios, in contrast, remain very stable—even during crises.

[26]This data stems from the Worldwide Governance Indicators (World Bank). Latin American economies cover almost the entire range of this index. This variation allows us to precisely identify the impact of institutional quality.

[27]This is the coefficient $LIBERALIZATION$.

[28]This is the sum of the coefficients $LIBERALIZATION$ and $CRISIS * LIBERALIZATION$, multiplied with 0.32 (the standard deviation of $LIBERALIZATION$).

Table 8: The Crisis Effects

	without a deposit insurance			with a deposit insurance and good institutions			with a deposit insurance and weak institutions		
	(1) STABLEFUNDING	(2) RETAIL	(3) INTERBANK	(4) STABLEFUNDING	(5) RETAIL	(6) INTERBANK	(7) STABLEFUNDING	(8) RETAIL	(9) INTERBANK
STABLEFUNDING (t-1)	0.500***			0.721***			0.709***		
	(7.81)			(18.71)			(13.88)		
RETAIL (t-1)		0.427***			0.773***			0.768***	
		(8.48)			(24.15)			(19.50)	
INTERBANK (t-1)			0.511***			0.676***			0.703***
			(10.04)			(17.73)			(13.17)
LIBERALIZATION	-7.510*	-2.803	23.820***	-0.083	1.088	0.175	-6.002**	7.420***	-1.072
	(-1.75)	(-0.40)	(3.03)	(-0.13)	(1.49)	(0.34)	(-2.42)	(3.26)	(-0.70)
CRISIS $\times$ LIBERALIZATION	11.841	-5.627	-39.440***	1.415	-1.035	-0.412	4.949*	-5.275**	-1.439
	(1.46)	(-0.64)	(-4.36)	(1.29)	(-0.86)	(-0.54)	(1.86)	(-2.48)	(-0.66)
CRISIS	-10.918	6.845	33.254***	0.168	-0.338	0.575	-4.745**	3.826**	1.079
	(-1.52)	(0.86)	(4.15)	(0.26)	(-0.42)	(1.06)	(-2.09)	(2.06)	(0.58)
Bank-Level Controls	YES	YES	YES	YES	YES	YES	YES	YES	YES
Macroeconomic Controls	YES	YES	YES	YES	YES	YES	YES	YES	YES
Obs	1355	2179	1338	3670	4530	4006	1621	2097	1653

In these regressions, we determine the effects of financial liberalization on the stable funding ratios, the shares of retail deposits and the fractions of interbank borrowing during financial crises. The regressions are estimated separately for countries without deposit insurance, with deposit insurance/ good institutions and with deposit insurance/weak institutions. The financial liberalization measure takes the values between 0 (the capital account closed off) and 1 (fully liberalized) and we also enable it to interact with a dummy that is one for countries with a local crisis or if there is a global banking sector crisis. We include the lagged dependent variables and bank-level and macro controls in the specifications. All of the Blundell-Bond tests are implemented with Stata's xtabond2 command (Roodman, 2009a), using 4 lags of the variables as instruments.The t-statistics are presented in parentheses and the standard errors are corrected by Windmeijer's (2005) correction procedure.

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

In this sense, this test also justifies our definition of the stable funding ratio, which excludes retail deposits: the latter appear to be generally less stable than equity and long-term funding, as depositors usually withdraw large proportions of their savings during crises, except for the handful of countries in the region that have relatively strong institutions or a deposit insurance scheme. Therefore, interpreting retail deposits as a stable type of funding (as it is typically done in analyses of banking systems in advanced economies) is contingent on a reliable deposit insurance scheme. Only in countries with good institutions, financial crises do not affect the access to retail funding.

To sum up, large banks are a potential risk for financial stability because international liberalization induces them to replace stable funding with interbank funding that is shown to be volatile across countries and across varying institutional qualities. But smaller banks may also be at risk, as they replace stable funding with retail deposits, which in Latin America have been quite volatile during financial crises and especially so in countries with less sound institutional frameworks. This evidence is—among other things—suggestive of the desirability of sounder deposit insurance schemes and better institutions overall. To the extent that these make customer deposits a more stable form of funding, they reduce the risk of bank runs once the country's capital account is liberalized.

6 Sensitivity Analysis

In this section, we show some final robustness checks. They include: (i) estimating our baseline model by using alternative measures for financial liberalization and (ii) splitting our sample into alternative sub-periods.

In the first test, we replace the capital inflow restriction measure with one that is applicable to net flows (gross capital inflows minus gross capital outflows). This helps establish the robustness of our results to the choice of external liberalization proxy (Table A.2, column (1) and (2)). Moreover, in columns (3) and (4), we dis-aggregate our liberalization index that measures the overall capital inflow restrictions and focus on the net inflow openness indicator applicable to financial instruments that benefit banks. In particular, we take the average of the net openness of money market flows, financial credit flows and bond flows and replace the liberalization index with such a composite indicator. Attendant results indicate that changing the liberalization index does not affect the economic or statistical significance once macroeconomic controls are in place. Except for the specification of column (3), when those controls are dropped, the coefficient on $LIBERALIZATION$ remains sizable and statistically significant.

In the second test, we replace $LIBERALIZATION$ with (i) the Chinn-Ito index (Chinn and Ito, 2006), which is also a de-jure measure of financial openness and (ii) the sum of external assets and liabilities relative to GDP ($EXTERNALWEALTH$), which is a de-facto measure of financial integration, proposed by Lane and Milesi-Ferretti (2007). As in the Schindler index of Fernández et al. (2015a), both of these indices take higher values in liberalized countries. The pairwise correlation between our index and the Chinn-Ito index and $EXTERNALWEALTH$ is equal to 78.6% and 35.7%, respectively, underlining that our measure of financial integration seems to be an appropriate measure of international financial openness. Table A.3 shows that the effect of liberalization on the funding decisions of banks remains consistent when we rely on other de-facto and de-jure measures of international financial integration. In particular, an increase in the Chinn-Ito index by one standard deviation (that is equal to 1.34) lowers the stable funding ratios on impact by 0.68-0.87 pp and a rise in $EXTERNALWEALTH$ by one standard deviation (that is equal to 75.25) reduces the proportion of equity and long-term funding relative to total assets by 1.05-1.28 pp in the short-run. Therefore, the economic effects of both indices are higher than our baseline estimates. Yet, as argued in Section 2.2, the Schindler index of Fernández et al. (2015a) allows us to focus on de-jure capital inflow restrictions, rather than on overall restrictions (of either a de-jure or a de-facto nature), which are likely to be less accurate metrics, in most circumstances, of the effects of liberalization on bank funding structures. Be that as it may, the results of Table A.2 suggest that our baseline estimates are—if anything—on the conservative side, further buttressing the thrust of this paper's overall finding.

Our final robustness test consists of estimating our baseline regression model over subsamples. Although we lose some variation in the liberalization measure, this adjustment might be important because the Bankscope database has a better coverage for the period 2000-2013.[29] As a consequence, in columns (1) and (2) of Table A.4, we drop the years

[29]Compare Table 1.

before 1997 and in columns (3) and (4), we drop the years before 2000. Moreover, in columns (5) and (6), we restrict the sample to commercial banks, cooperative banks and savings banks and, hence, we drop, e.g., governmental institutions from our dataset. All of these adjustments do not change the coefficients in our regressions significantly. This makes us confident that the varying number and the different types of banks in our sample do not bias our results.

7 Concluding Remarks

Previous research focused on the effects of international financial integration and cross-border capital flows on the asset side of banks. The effects on the funding structure of banks have remained underexplored. This paper sought to fill some of that gap.

We find that stronger affiliation to the international capital markets is associated with a reduction in the stable funding ratios of banks, i.e., less equity and less long-term funding. We additionally identify an amplified effect in banks with more opaque balance sheets, suggesting that higher asymmetries among local banks and international lenders tend to lower stable funding ratios. The importance of informational asymmetries is also corroborated by evidence that the effect of financial integration is overproportional in banks whose lenders are geographically more distant. This result highlights one potential advantage of a higher presence of foreign financial institutions in developing countries.

Our results also show that small and large banks are affected differently by international financial liberalization, providing further support for the role of information asymmetries in the dynamics of banks' funding structures after financial integration, as well as the important role of the heterogeneity of institutional quality within the same developing region in shaping these dynamics.

These findings complement and provide further support to some of the current wisdom on how international financial integration can increase the propensity for financial instability. All else constant, the average bank will tend to reduce the share of stable funding and increase its reliance on short-term funding with financial integration. This result suggests that financial institutions are more prone to rollover risks following post-liberalization bursts of capital inflows and large current account deficits. Greater distance from main lending countries and weaker domestic institutions exacerbate such an effect.

These findings should not to be interpreted as a rejection of the many benefits from international financial integration, but rather highlight the importance of macroprudential regulations as countries become financially more open. Our results also indicate that macroprudential policies have a greater role in countries that are more remote from world financial centers and where the quality of institutions rank lower on a global scale.

References

[1] Altunbas, Y., L. Gambacorta, and D. Marques-Ibanez: 2014, 'Does Monetary Policy Affect Bank Risk?'. *International Journal of Central Banking* **10**(1), 95–136.

[2] Anderson, T. W. and C. Hsiao: 1981, 'Estimation of Dynamic Models with Error Components'. *Journal of the American Statistical Association* **76**(375), 598–606.

[3] Arellano, M. and S. Bond: 1991, 'Some Tests of Specification for Panel Data: Monte Carlo Evidence and an Application to Employment Equations'. *Review of Economic Studies* **58**(2), 277–297.

[4] Arellano, M. and O. Bover: 1995, 'Another Look at Instrumental Variables Estimation of Error-Component Models'. *Journal of Econometrics* **68**(1), 29–51.

[5] Babin, A.: 2015, 'Home-Host Banking Issues and Non-Core Funding—Evidence from Central and Eastern Europe'. *Open Economics Review* **26**(3), 447–477.

[6] Barth, J. R., G. Caprio Jr., and R. Levine: 2001, 'The Regulation and Supervision of Banks Around the World: A New Database'. *University of Minnesota Financial Studies Working Paper* **0006**.

[7] Bednarek, P., V. Dinger, and N. von Westernhagen: 2015, 'Fundamentals Matter: Idiosyncratic Shocks and Interbank Relations'. *Bundesbank Discussion Paper* **44/2015**.

[8] Bekaert, G. and C. R. Harvey: 2000, 'Foreign Speculators and Emerging Equity Markets'. *Journal of Finance* **55**(2), 565–613.

[9] Berger, A. N. and C. H. S. Bouwman: 2009, 'Bank Liquidity Creation'. *Review of Financial Studies* **22**(9), 3779–3837.

[10] Blundell, R. and S. Bond: 1998, 'Initial Conditions and Moment Restrictions in Dynamic Panel Data Models'. *Journal of Econometrics* **87**(1), 115–143.

[11] Booth, L., D.-K. A. Aivazian, V., and V. Maksimovic: 2001, 'Capital Structures in Developing Countries'. *Journal of Finance* **56**(1), 87–130.

[12] Boyd, J. H. and D. E. Runkle: 1993, 'Size and Performance of Banking Firms: Testing the Predictions of Theory'. *Journal of Monetary Economics* **31**(1), 47–67.

[13] Broner, F. A., G. Lorenzoni, and S. L. Schmukler: 2013, 'Why Do Emerging Economies Borrow Short Term?'. *Journal of the European Economic Association* **11**(1), 67–100.

[14] Brooks, S. M.: 2004, 'Explaining Capital Account Liberalization in Latin America: A Transitional Cost Approach'. *World Politics* **56**(3), 389–430.

[15] Bruno, V. and H. S. Shin: 2015, 'Cross-Border Banking and Global Liquidity'. *Review of Economic Studies* **82**(2), 535–564.

[16] Calomiris, C. W. and C. M. Kahn: 1991, 'The Role of Demandable Debt in Structuring Optimal Banking Arrangements'. *American Economic Review* **81**(3), 497–513.

[17] Calvo, G. A. and C. A. Végh: 1999, *Inflation Stabilization and BOP Crises in Developing Countries*. eds: John Taylor and Michael Woodford, North Holland. in: Handbook of Macroeconomics.

[18] Catão, L. A. V. and G. M. Milesi-Ferretti: 2014, 'External Liabilities and Crises'. *Journal of International Economics* **94**(1), 18–32.

[19] Cerutti, E., S. Claessens, and D. Puy: 2015, 'Push Factors and Capital Flows to Emerging Markets: Why Knowing Your Lender Matters More Than Fundamentals'. *IMF Working Paper* **15/127**.

[20] Cetorelli, N. and L. S. Goldberg: 2012, 'Banking Globalization and Monetary Transmission'. *Journal of Finance* **67**(5), 1811–1843.

[21] Chinn, M. D. and H. Ito: 2006, 'What Matters For Financial Development? Capital Controls, Institutions, and Interactions'. *Journal of Development Economics* **81**(1), 163–192.

[22] Correa, R., T. Paligorova, H. Sapriza, and A. Zlate: 2015, 'Cross-Border Bank Flows and Monetary Policy'. mimeo, Federal Reserve Board.

[23] Craig, B. R. and V. Dinger: 2014, 'Volatile Lending and Bank Wholesale Funding'. *Federal Reserve Bank of Cleveland Working Paper* **1417**.

[24] Cubillas, E. and F. González: 2014, 'Financial Liberalization and Bank Risk-Taking: International Evidence'. *Journal of Financial Stability* **11**(1), 32–48.

[25] Darvas, Z.: 2012a, 'Real Effective Exchange Rates for 178 Countries: A New Database'. *Bruegel Working Paper* **2012/06**.

[26] Darvas, Z.: 2012b, 'Compositional Effects on Productivity, Labour Cost and Export Adjustment'. *Bruegel Policy Contribution* **2012/11**.

[27] Darvas, Z.: 2012c, 'Compositional Effects on Productivity, Labour Cost and Export Adjustment: Detailed Results for 24 EU Countries'. *Bruegel Working Paper* **2012/11**.

[28] De Haas, R. and N. Van Horen: 2013, 'Running for the Exit? International Bank Lending During a Financial Crisis'. *Review of Financial Studies* **26**(1), 244–285.

[29] DeYoung, R. and T. Rice: 2004, 'How Do Banks Make Money? The Fallacies of Fee Income'. *Economic Perspectives* **Q4**, 34–51.

[30] Dinger, V. and D. M. te Kaat: 2016, 'Current Account Imbalances and Bank Risk-Taking'. mimeo, Osnabrück University.

[31] Dinger, V. and J. von Hagen: 2009, 'Does Interbank Borrowing Reduce Bank Risk?'. *Journal of Money, Credit and Banking* **41**(2-3), 491–506.

[32] European Central Bank: 2015, 'Financial Stability Review, November 2015'. Frankfurt a. M.

[33] Faulkender, M., M. J. Flannery, K. W. Hankins, and J. M. Smith: 2012, 'Cash Flows and Leverage Adjustments'. *Journal of Financial Economics* **103**(3), 632–646.

[34] Fernández, A., M. W. Klein, A. Rebucci, M. Schindler, and M. Uribe: 2015a, 'Capital Control Measures: A New Dataset'. *NBER Working Paper* **20970**.

[35] Fernández, A., A. Rebucci, and M. Uribe: 2015b, 'Are Capital Controls Countercyclical?'. *Journal of Monetary Economics* **76**, 1–14.

[36] Flannery, M. J.: 1986, 'Asymmetric Information and Risky Debt Maturity Choice'. *Journal of Finance* **41**(1), 19–37.

[37] Flannery, M. J. and K. W. Hankins: 2013, 'Estimating Dynamic Panel Models in Corporate Finance'. *Journal of Corporate Finance* **19**, 1–19.

[38] George, C.: 2015, 'Bank Size, Leverage, and Financial Downturns'. *FDIC Center for Financial Research Paper* **2015-01**.

[39] Gourinchas, P.-O. and M. Obstfeld: 2012, 'Stories of the Twentieth Century for the Twenty-First'. *American Economic Journal: Macroeconomics* **4**(1), 226–265.

[40] Gropp, R. and F. Heider: 2010, 'The Determinants of Bank Capital Structure'. *Review of Finance* **14**(4), 587–622.

[41] Hahm, J.-H., H. S. Shin, and K. Shin: 2013, 'Noncore Bank Liabilities and Financial Vulnerability'. *Journal of Money, Credit and Banking* **45**(8), 3–36.

[42] Hauswald, R. and R. Marquez: 2006, 'Competition and Strategic Information Acquisition in Credit Markets'. *Review of Financial Studies* **19**(3), 967–1000.

[43] Hoffmann, M. and I. Stewen: 2016, 'Holes in the Dike: The Global Savings Glut, U.S. House Prices and the Long Shadow of Banking Deregulation'. *CAMA Working Paper* **6/2016**.

[44] Innes, R. D.: 1990, 'Limited Liability and Incentive Contracting with Ex-Ante Action Choices'. *Journal of Economic Theory* **52**(1), 45–67.

[45] Ioannidou, V. P., S. Ongena, and J.-L. Peydró: 2015, 'Monetary Policy, Risk-Taking, and Pricing: Evidence from a Quasi-Natural Experiment'. *Review of Finance* **19**(1), 95–144.

[46] Jeanne, O.: 2009, 'Debt Maturity and the International Financial Architecture'. *American Economic Review* **99**(5), 2135–2148.

[47] Johnson, S. A.: 2015, 'Debt Maturity and the Effects of Growth Opportunities and Liquidity Risk on Leverage'. *Review of Financial Studies* **16**(1), 209–236.

[48] Jordà, Ò., M. Schularick, and A. M. Taylor: 2013, 'When Credit Bites Back: Leverage, Business Cycles, and Crises'. *Journal of Money, Credit, and Banking* **45**(2), 3–28.

[49] Jungherr, J.: 2016, 'Bank Opacity and Financial Crises'. *Barcelona GSE Working Paper* **882**.

[50] Kang, J.-K. and R. M. Stulz: 1997, 'Why is there a Home Bias? An Analysis of Foreign Portfolio Equity Ownership in Japan'. *Journal of Financial Economics* **46**(1), 3–28.

[51] Laeven, L. and F. Valencia: 2013, 'Systemic Banking Crises Database'. *IMF Economic Review* **61**(2), 225–270.

[52] Lane, P. R. and G. M. Milesi-Ferretti: 2007, 'The External Wealth of Nations Mark II: Revised and Extended Estimates of Foreign Assets and Liabilities, 1970-2004'. *Journal of International Economics* **73**(2), 223–250.

[53] Lee, C.: 2011, 'Firms' Capital Structure Decisions and Product Market Competition: A Theoretical Approach'. *Research in Business and Economics Journal* **3**.

[54] Lucey, B. M. and Q. Zhang: 2011, 'Financial Integration and Emerging Markets Capital Structure'. *Journal of Banking and Finance* **5**(35), 1228–1238.

[55] Mendoza, E. G. and M. E. Terrones: 2012, 'An Anatomy of Credit Booms and their Demise'. *Economia Chilena* **15**(2), 4–32.

[56] Mian, A.: 2006, 'Distance Constraints: The Limits of Foreign Lending in Poor Economies'. *Journal of Finance* **61**(3), 1465–1505.

[57] Morais, B., J.-L. Peydró, and C. Ruiz: 2015, 'The International Bank Lending Channel of Monetary Policy Rates and Quantitative Easing: Credit Supply, Reach-for-Yield, and Real Effects'. *World Bank Policy Research Working Paper* **7216**.

[58] Myers, S. C. and N. S. Majluf: 1984, 'Corporate Financing and Investment Decisions When Firms Have Information That Investors Do Not Have'. *Journal of Financial Economics* **13**(2), 187–221.

[59] Nickel, S.: 1981, 'Biases in Dynamic Models with Fixed Effects'. *Econometrica* **49**(6), 1417–1426.

[60] Ongena, S., J.-L. Peydró, and N. van Horen: 2015, 'Shocks Abroad, Pain at Home? Bank-Firm Level Evidence on the International Transmission of Financial Shocks'. *IMF Economic Review* **63**(4), 698–750.

[61] Popov, A. and G. F. Udell: 2012, 'Cross-Border Banking, Credit Access, and the Financial Crisis'. *Journal of International Economics* **87**(1), 147–161.

[62] Reinhart, C. M. and K. S. Rogoff: 2008, 'Is the 2007 US Sub- prime Financial Crisis So Different? An International Historical Comparison'. *American Economic Review* **98**(2), 339–344.

[63] Rodrik, D. and A. Velasco: 1999, 'Short-Term Capital Flows'. *NBER Working Paper* **7364**.

[64] Roodman, D.: 2009a, 'How to do Xtabond2: An Introduction to Difference and System GMM in Stata'. *Stata Journal* **9**(1), 86–136.

[65] Roodman, D.: 2009b, 'A Note on the Theme of Too Many Instruments'. *Oxford Bulletin of Economics and Statistics* **71**(1), 135–158.

[66] Sarig, O. H.: 1998, 'The Effect of Leverage on Bargaining with a Corporation'. *Financial Review* **33**(1), 1–16.

[67] Song, F. and A. V. Thakor: 2007, 'Relationship Banking, Fragility, and the Asset-Liability Matching Problem'. *Review of Financial Studies* **20**(6), 2129–2177.

[68] Stiglitz, J. E.: 2000, 'Capital Market Liberalization, Economic Growth, and Instability'. *World Development* **28**(6), 1075–1086.

[69] Stohs, M. H. and D. C. Mauer: 1996, 'The Determinants of Corporate Debt Maturity Structure'. *Journal of Business* **69**(3), 279–312.

[70] Stulz, R. M.: 1999, ‘International Portfolio Flows and Security Markets’. *Working Paper* **99-3**.

[71] Taylor, A. M.: 2015, ‘Credit, Financial Stability, and the Macroeconomy’. *Annual Review of Economics* **7**(1), 309–339.

[72] Temesvary, J., S. Ongena, and A. L. Owen: 2015, ‘A Global Lending Channel Unplugged? Does U.S. Monetary Policy Affect Cross-Border and Affiliate Lending by Global U.S. Banks?’. *CFS Working Paper* **511**.

[73] Trubek, D. M., H. A. Garcia, D. R. Coutinho, and A. Santos: 2013, *Law and the New Developmental State: The Brazilian Experience in Latin American Context*. Cambridge University Press, Cambridge.

[74] Windmeijer, F.: 2005, ‘A Finite Sample Correction for the Variance of Linear Efficient Two-Step GMM Estimators’. *Journal of Econometrics* **126**(1), 25–51.

[75] Wu, J. C. and F. D. Xia: 2016, ‘Measuring the Macroeconomic Impact of Monetary Policy at the Zero Lower Bound’. *Journal of Money, Credit, and Banking* **48**(2–3), 253–291.

Table A.1: Description of the Variables

Variable	Description	Unit	Source
STABLEFUNDING	(equity + long-term funding) / total assets	%	Bankscope, own calculations
RETAIL	customer deposits / total assets	%	Bankscope, own calculations
INTERBANK	interbank liabilities / total assets	%	Bankscope, own calculations
SIZE	ln (total assets)	ln (million x)	Bankscope, own calculations
LIQUIDITY	liquid assets / total assets	%	Bankscope, own calculations
PROFITABILITY	net income / total assets	%	Bankscope, own calculations
NONINTERESTINCOME	non-interest income / gross revenues	%	Bankscope, own calculations
RISK	(impaired loans - reserves for impaired loans) / equity	%	Bankscope, own calculations
DUMMY-LARGE	Dummy=1 if assets above 95th percentile of annual distribution, 0 below 90th percentile	0/1	Bankscope, own calculations
LIBERALIZATION	(1 - Schindler inflow restrictions index)	-	Fernández et al. (2015a), own calculations
PERCAPITAGDP	PPP adjusted per capita GDP	x/1000	WEO, own calculations[a]
CAPITALREQUIREMENT	The regulatory minimum capital requirement	%	Barth et al. (2001), own calculations
INFLATION	The relative change in the CPI index	%	WEO, own calculations
REALEXCHANGERATE	Real effective exchange rate (based on CPI)	-	Darvas (2012a, 2012b, 2012c)
DISTANCE	The weighted distance between borrower and lender countries, details in equation (2)	-	BIS, own calculations[b]
CRISIS	Dummy = 1 if there is a local or global banking sector crisis	0/1	Laeven and Valencia (2013), own calculations

[a] World Economic Outlook Database, IMF.
[b] Consolidated Banking Statistics.

Table A.2: Robustness Test (1)

	(1) STABLEFUNDING	(2) STABLEFUNDING	(3) STABLEFUNDING	(4) STABLEFUNDING
STABLEFUNDING (t-1)	0.661***	0.669***	0.664***	0.675***
	(14.60)	(16.77)	(14.14)	(17.51)
LIBERALIZATION	-2.879**	-3.161***	-0.962	-2.409***
	(-2.46)	(-3.56)	(-1.06)	(-3.38)
Bank-Level Controls	YES	YES	YES	YES
Macroeconomic Controls	NO	YES	NO	YES
Obs	6660	6646	6660	6646

In this first robustness check, we replace the overall measure for inflow restrictions with (i) net capital flow restrictions (columns 1, 2), and (ii) the net flow restrictions on money market, bond and credit flows (columns 3,4). These variables in liberalized countries are equal to 1.The dependent variable is the sum of equity and long -term funding (> 1 year) relative to total assets. We include the lagged dependent variables and bank-level and macroeconomic controls in some specifications. All of the Blundell-Bond tests are estimated with Stata's xtabond2 command (Roodman, 2009a), using 4 lags of the variables as instruments. The t- statistics are shown in parentheses, using Windmeijer (2005) corrected standard errors.

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

Table A.3: Robustness Test (2)

	(1) STABLEFUNDING	(2) STABLEFUNDING	(3) STABLEFUNDING	(4) STABLEFUNDING
STABLEFUNDING (t-1)	0.661***	0.678***	0.672***	0.657***
	(14.79)	(17.03)	(14.09)	(15.10)
CHINNITO	-0.650***	-0.504***		
	(-4.01)	(-3.64)		
EXTERNALWEALTH			-0.014***	-0.017***
			(-4.03)	(-4.42)
Bank-Level Controls	YES	YES	YES	YES
Macroeconomic Controls	NO	YES	NO	YES
Obs	6660	6646	5609	5595

In this next robustness check, we replace the overall measure for inflow restrictions with (i) the Chinn- Ito index (Chinn and Ito, 2006) and (ii)the sum of external assets and liabilities (Lane and Milesi-Ferreti, 2007). The dependent variable is the sum of equity and long-term funding(>1 year) relative to total assets. We include the lagged dependent variables and bank - level and macroeconomic controls in some specifications. All of the Blundell-Bond tests are estimated with Stata's xtabond2 command (Roodman, 2009a), using 4 lags of the variables as instruments.The t-statistics are presented in parentheses and the standard errors are corrected by Windmeijer (2005).

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

Table A.4: Robustness Test (3)

	(1) STABLEFUNDING	(2) STABLEFUNDING	(3) STABLEFUNDING	(4) STABLEFUNDING	(5) STABLEFUNDING	(6) STABLEFUNDING
STABLEFUNDING (t-1)	0.664^{***}	0.681^{***}	0.682^{***}	0.685^{***}	0.664^{***}	0.672^{***}
	(14.60)	(17.00)	(13.30)	(14.83)	(16.34)	(19.68)
LIBERALIZATION	-1.350	-1.552^{**}	-1.429	-1.794^{**}	-1.271^{*}	-1.293^{**}
	(-1.61)	(-2.41)	(-1.60)	(-2.47)	(-1.73)	(-2.23)
Bank-Level Controls	YES	YES	YES	YES	YES	YES
Macroeconomic Controls	NO	YES	NO	YES	NO	YES
Obs	6445	6435	5505	5505	4976	4962

In this robustness check, we estimate equation (1) by restricting the sample period to 1997-2013 (columns 1,2), by restricting the sample period to 2000-2013 (columns 3,4) and by keeping only commercial, cooperative as well as savings banks (columns 5,6). The dependent variable is the sum of equity and long-term funding over total assets. The financial liberalization measure takes the values between 0 (capital account closed off) and 1 (fully liberalized). We include the lagged dependent variable and bank- level and macroeconomic controls in some specifications. We estimate the regressions with Stata's xtabond2 command (Roodman, 2009a), using 4 lags of the variables as instruments. The t-statistics are shown in parentheses, calculated with standard errors that are corrected by Windmeijer's (2005) correction procedure.

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