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Capital account liberalization and the composition of bank liabilities

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Using a sample of almost 600 banks in Latin America, we show that capital account liberalization lowers the share of equity and raises the share of interbank funding in total liabilities of the banking system. These shifts are mostly due to large banks; smaller banks, instead, increase their resort to retail funding by offering higher average deposit interest rates than larger banks. We also find significant differences in the behavior of banks with seemingly greater information opacity. These findings have positive implications for macro-prudential regulation.

1. Introduction

Lower controls on a country's capital account can increase the conditional probability of macro-financial crises by facilitating the accumulation of foreign liabilities (Reinhart and Rogoff, 2008; Gourinchas and Obstfeld, 2012; Claessens and Ghosh, 2013; Catão and Milesi-Ferretti, 2014; Claessens, 2017).[1] In examining the channels through which abundant global liquidity and capital flows raise crisis risk, many papers have looked at the role of bank lending to private firms and the government—flows that lie squarely on the asset side of banks' balance sheets (Popov and Udell, 2012; Jordà et al., 2013; Lane and McQuade, 2014; Taylor, 2015; Ongena et al., 2015; Correa et al., 2015; Baskaya et al., 2017; Temesvary et al., 2018; Morais et al., 2019; Dinger and te Kaat, 2020; Hoffmann and Stewen, 2020; te Kaat, forthcoming). On the liability side, while there has been work on how bank leverage responds to swings in global liquidity and risk aversion (Bruno and Shin, 2015a,b), and on how capital control regulations affect the cost and volume of international borrowing and lending to firms (Bonfiglioli, 2008; Beakart et al., 2011; Varela, 2018; Ahnert et al. 2021), little attention has been devoted to how the distinct components of banks' liabilities shift in response to changes in capital controls and how those shifts are conditioned by bank-specific characteristics (large vs. small, foreign vs. domestically-owned, having a more vs. a less opaque balance sheet).

This paper aims to fill some of this gap in the literature. We examine the response of the various liability components of banks—namely, equity, retail deposits, interbank deposits, bonds, other short-term debt, and non-interest liabilities—to

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[1] In this paper, the capital account encompasses what the IMF in its Balance of Payments and International Investment Position Manual (sixth edition) calls financial account.

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changes in capital control regulations and ask whether and how such responses differ across smaller and larger banks, domestic vs. foreign-owned, and those with seemingly more vs. less opaque balance sheets. We do so using bank-level data from Bureau van Dijk's Bankscope database for 17 Latin American countries over 1995–2013. Focusing on Latin America during this period is particularly suitable for the purpose of this investigation because of the region's pattern of liberalization in external capital accounts, which was not only far-reaching, but also displayed considerable cross-country heterogeneity through the 1990s, 2000s and early 2010s, aiding identification.[2] We further aid identification by controlling our estimates for a variety of country-specific macro and global financial variables, using a more accurate metric of capital controls developed in Fernández et al. (2016), and allowing for some possible endogeneity of capital controls. The sizable dimension of our panel data—across countries, banks, and time—coupled with our use of a rich set of macro-financial covariates in turn allows our findings to speak to the intersection of the prominent literature on banks, financial globalization and macro-financial risk.

A clear conceptual motivation for our study stems from the expected effects of capital controls on the relative costs of distinct funding sources to banks, as well as on the nature of foreign investors' participation in domestic banks. For instance, resort to cheaper foreign borrowing resulting from lower taxes (or more friendly regulation) on foreign borrowing should be expected to crowd-out resort to some traditional sources of domestic borrowing (such as retail deposits) and may thus shift the composition of bank liabilities to foreign interbank and/or international bond issuance. Likewise, cheaper and (perceived) less constrained resort to foreign funding may in turn affect banks' incentives to build capital buffers. No less importantly perhaps, the prototype of investor and his/her information set may also be importantly affected by capital account regulations. Assume, for instance, that foreign investors may still face, after liberalization, more costly or less available information on the underlying fundamentals and lending strategies of a domestic bank relative to what national investors face. If so, once foreign investors increase their participation in banks' funding markets, the degree of asymmetric information between banks and their median investors should therefore rise. Such a rise in asymmetric information lowers the cost of funding sources that are less sensitive to information asymmetries, such as short-term debt (including through foreign interbank markets), relative to long-term debt and equity (Myers and Majluf, 1984), thereby tilting banks' incentives toward raising the share of short-term liability (including through the interbank market) in total liabilities.[3] At the same time, it should also be expected that banks that are foreign-owned, larger, and with less opaque balance sheets experience a stronger push factor with capital account liberalization. This in turn will change the liability structure of the consolidated bank system toward those of larger and foreign banks, which are also sometimes less capitalized and often have far more extensive links to foreign counterpart institutions—with attendant implications for macro-financial risk. Overall, there are thus substantive reasons to expect that changes in capital account regulations or controls will affect the percentage shares of banks' main liability components and that this may have broader macro implications.

Our findings are as follows. First, we find that capital account openness is associated with lower capital-to-asset ratios and increases in banks' interbank liabilities. All other liability side variables are mostly unaffected. In economic terms, a one-standard deviation increase in capital account openness is associated with 0.38 percentage point ("pp" henceforth) reductions in banks' equity ratios and 0.43 pp higher interbank funding ratios in the short-run. The respective long-run effects are 1.6 pp (for capital-to-asset ratios) and 2.8 pp (for interbank liabilities). The economic significance of these results has been fleshed out in previous work: according to the ECB (2015), for instance, a one-pp decrease in the Tier 1 capital ratios raises the odds ratio (the probability of distress relative to non-distress) by 35–39% (see also Altunbas et al., 2014). Second, we find that the economic and statistical significance of these effects is dominated by periods of high real domestic money market interest rate spreads relative to the world's main financial center—the US. Specifically, the documented shifts in the liability composition of banks are largely a preserve of capital account liberalization measures enacted during periods of low US interest rates. Third, some key results also arise from the interaction of capital account openness with bank size, foreign ownership, and an indicator of balance sheet opacity (the ratio of impaired loans relative to equity, predicated under the assumption that banks with lower credit risks typically have a less opaque balance sheet). We find that especially larger and informationally less opaque banks raise their interbank liabilities and lower their capital-to-asset ratios disproportionately more in response to capital account liberalization. While this possibly reflects higher regulatory margins in the pre-liberalization state and/or larger banks' greater latitude to operate with lower capital ratios, such a post-liberalization shift also appears to reflect wider access to cheaper interbank funding, which motivates a substitution of less risky (capital) by higher risk funding (interbank borrowing). In contrast, smaller banks increase their reliance on retail deposits in the wake of capital account liberalizations by offering higher interest rates on deposits, leading to the migration of deposits to smaller banks. Either way, the overall effect is in the direction of increasing systemic risk in the wake of liberalization, all else constant.

As alluded to earlier, these findings speak to the broader literature relating capital account regulations and international capital flows to banks' funding and macro-financial risk. As in Bruno and Shin (2015a,b), we show that lower foreign borrow-

[2] 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. (2019), which would call for more evolved and (arguably) less consensual model restrictions to help identification of regional factors.

[3] Indeed, evidence from the behavior of broad stock price indices and bond spreads following major capital account liberalizations is consistent with this conjecture (Stulz, 1999; Bekaert and Harvey, 2000), as is the evidence that short-term debt flows are the dominant type of cross-border capital flows to emerging market economies (e.g., Henry, 2007; Kose et al., 2009). Studies in non-bank corporate finance also find empirical support for a shift towards short-term debt due to informational frictions that change the relative costs of funding (Stohs and Mauer, 1996; Johnson, 2015).

ing cost— in this case due to capital account liberalization—tend to raise leverage, particularly during periods of low global interest rates. Our paper goes beyond their findings, however, by also documenting effects across distinct tiers of bank liabilities and across bank size, foreign ownership and balance sheet opaqueness—dimensions which are all very relevant for systemic risk assessment. The result on the greater importance of larger banks in heightening aggregate financial risk is also broadly in line with the findings of Baskaya et al. (2017), who show that higher credit growth in Turkey is mostly driven by bursts of foreign capital inflows channeled through larger banks, responding to a supply side capital push external to the country. In contrast, we show that smaller banks increase their reliance on retail deposits in the wake of capital account liberalizations, leading to the migration of deposits to smaller banks. These in turn are well-know to be more susceptible to bank runs and flight-to-safety once macro-financial distress kicks in.

Our results also broaden the findings of a literature on the determinants of banks' funding decisions in general which has glossed over how such decisions are affected by capital controls (Song and Thakor, 2007; Berger and Bouwman, 2009; Dinger and von Hagen, 2009; Hahm et al., 2013; Craig and Dinger, 2014). Our results on the rising share of (short-term) interbank funding, higher leverage, and seemingly heightened asymmetric information sensitivity also draws attention to the financial risk dimension that permeates numerous works on the effects of capital account liberalization/capital controls on the real economy (Henry, 2003; Voth, 2003; Henry, 2007; Kose et al., 2009; Levchenko et al., 2009; Larrain and Stumpner, 2017), and on the relationship between external financial openness and financial risk (Daniel and Jones, 2007; Martinez-Miera and Repullo, 2017). In particular, and consistent with the recent study by Ahnert et al. (2020) on the impact of macro-prudential foreign exchange regulations on currency mismatch and lending risk, our results offer a cautioning tale to the positive effects of financial liberalization on the productivity of non-financial firms (Bonfiglioli, 2008; Bekaert et al., 2011; Lucey and Zhang, 2011; Agca et al., 2015; Varela, 2018).

The remainder of this paper is structured as follows. Section 2 describes the institutional setting and trends in capital account liberalization in Latin America. Section 3 presents the data set and summary statistics. Section 4 lays out the econometric methodology and reports our baseline results. In Section 5, we test whether our baseline results are amplified by less opaque banks. The effects on smaller banks are investigated in Section 6. Section 7 performs various robustness checks. Section 8 concludes the paper.

2. Background Facts

Fig. 1 displays the average degree of capital account openness over the period of 1980–2013, and the corresponding one-sd bands around the mean.

The reduction in capital controls in Latin America trended up between the early 1990s through 2007, and has been partly reversed since the onset of the global financial crisis. The wide standard deviation bands also indicate that there is significant cross-country variation in external financial openness. This contrasts with the experience of other emerging market regions of Asia and Central and Eastern Europe, where the cross-country variation was about one-half lower.[4]

In much of the region, the trend towards greater external financial liberalization has been 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. In countries with IMF programs, those were an additional prodding force. Another determinant was a global trend towards external 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 that trade treaty, the decision to join NAFTA was instrumental to the disbanding of the stringent system of capital controls. Elsewhere in the region, other national-specific elements also played a role as, for instance, Brazil in the early 1990s under president Collor de Mello, when trade and capital flows were liberalized as a political response to the inefficiency of domestic monopolies in manufacturing and finance.[5]

These considerations suggest that capital account restrictions have an exogenous component relative to the funding structure of banks. Yet, it has been shown that capital controls often react to macroeconomic and financial variables, including capital inflows, foreign exchange reserves and fiscal balances (Cardoso and Goldfajn, 1998; Aizenman and Pasricha, 2013; Magud et al., 2011), even if capital controls appear to be broadly a-cyclical in a broad cross-section of countries (Fernández et al., 2015). Moreover, theory indicates that financial stability, which may be affected by the liability structure of banks, is a key motivation for capital control regulations, even if some studies fail to identify a systematic policy response of capital controls to macro-financial developments (Erten et al., forthcoming). Finally, and particularly in the case of Latin America where bank concentration has been high and the political influence of large banks non-trivial, there is also the possibility that bank funding structures and bankers' behavior might affect the timing, the nature and/or the strength of capital control regulations. Thus, one cannot readily discard the possibility that capital control decisions also have a potentially endogenous component related to the liability structure indicators considered in this paper.

[4] For further break-downs of the index by region, sub-indices and sub-periods, see Fernández et al. (2016).

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

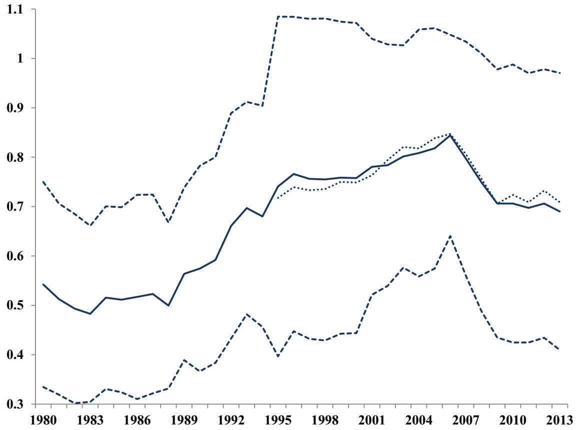

Fig. 1. The blue solid line displays the average degree of capital account openness (proxied by the overall Schindler index from Fernández et al., 2016) in Latin America over the 1980–2013 period, using the Quinn (1997) index to extrapolate it backwards until 1980. The dashed lines are the corresponding one-sd bands around the mean. The dotted line depicts the inflow-only component of the Schindler index.

While recognizing that assuming full exogeneity of capital controls is problematic in a cross-section of countries, our analysis guards against the possibility that partial endogeneity might bias our results in three distinct ways: first, by relying on an identification strategy hinging on the heterogeneity of banks at the micro-level and using a broad national cross-section of banks to this effect; second, by using GMM estimators that rely on lags of the right-hand side variables; and third, by presenting specifications in which we employ a government's partisanship indicator and an IMF program dummy as exogenous instruments. Specifics of each of these estimation procedures are discussed in Sections 4.1 and 7 below.

3. Data

3.1. Bank-Level Data

Our annual bank-level data spans the 1995–2013 period and the following 17 Latin American countries: Argentina, Bolivia, Brazil, Chile, Colombia, Costa Rica, Dominican Republic, Ecuador, El Salvador, Guatemala, Mexico, Nicaragua, Panama,[6] Paraguay, Peru, Uruguay and Venezuela.[7]

The variables are constructed from information provided in Bureau van Dijk's Bankscope database. We mostly include unconsolidated balance sheet data because consolidated statements might be affected by foreign subsidiaries.[8] Most banks report their balance sheet numbers in December of the respective year. For few banks that report those characteristics in the first five months of a year, we define them to belong to the previous year. After transforming all data into USD, implementing some data cleaning with regard to mergers by dropping bank observations where asset growth is larger than 100% or lower than −50% and after dropping implausible observations (e.g., negative assets, equity or loans), we obtain a sample of 8,278 bank-year observations.

Table 1 presents the number of financial institutions over time. Bankscope coverage is lower for the 1990s relative to the 2000s, which results in a lower number of banks in our data set for 1995–1999. As we will show in the sensitivity analysis presented in Section 7, we obtain qualitatively similar results for time sub-periods with a relatively constant number of banks. Further, Table 2 shows that most banks in our sample are located in the three largest economies of Latin America—Argentina, Brazil and Mexico.

We use this rich bank-level data set to break down bank liabilities into equity (*CAPITAL*), retail deposits (*DEPOSITS*), interbank funding (*INTERBANK*), other short-term debt (*OTHER SHORT − TERM DEBT*),[9] bonds (*BONDS*)[10] and non-interest bearing liabilities (*NON − INTEREST FUNDS*), all expressed as ratios relative to total assets.

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

[7] Three Latin American countries (Cuba, Honduras and Puerto Rico) are not covered because of missing data on their degree of external financial openness. We start our sample period in 1995 because both our bank-level data and the measure of capital account openness (the de-jure index of Fernández et al., 2016) are not available before.

[8] When banks only report consolidated statements, we include these in our regressions to increase the number of observations. As a bank that reports both consolidated and unconsolidated statements may have different bank names in Bankscope, we use Stata's reclink command, a module to probabilistically match records (Blasnik, 2010), with a minimum matching reliability of 0.99, to determine whether two statements belong to one bank. Only consolidated statements that have not been probabilistically matched to an unconsolidated one are included in our sample.

[9] This variable includes all short-term liabilities that are not interbank or retail deposits. For instance, it includes money market funds and corporate deposits.

[10] This variable basically includes all traded liabilities. However, long-term bonds with a share of 92% in all traded liabilities, are by far the most critical component.

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

	1995	2000	2005	2010	2013
number of banks	123	463	399	589	569

Table 2
The Distribution of Banks in our Sample across Countries.

Country	number of banks
Argentina	133
Bolivia	22
Brazil	220
Chile	79
Colombia	115
Costa Rica	111
Dom. Republic	35
Ecuador	68
El Salvador	44
Guatemala	51
Mexico	177
Nicaragua	26
Panama	76
Paraguay	33
Peru	46
Uruguay	38
Venezuela	91

Our bank-level data set further contains various explanatory variables that are likely to affect banks' funding structures. These include bank size (*SIZE*), defined as the logarithm of total assets, the ratio of impaired loans less reserves for impaired loans over total equity as a proxy for bank risk (*RISK*) and the share of non-interest income over gross revenues (*NONINTERESTINCOME*).

3.2. Macroeconomic Data

Our main regressor is the degree of capital account openness, proxied by the Schindler inflow index (Fernández et al., 2016). It is a new de-jure index of external 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 between zero (fully liberalized) and one. In our model, *LIBERALIZATION* is calculated as (1-Schindler inflow index) because—due to this transformation—higher values represent external financial liberalization, facilitating the interpretation of our results. One key advantage of this index is that it reports the openness of capital in- and outflows separately. For the analysis of this paper, focusing on inflow restrictions is beneficial because inflows of foreign capital are likely to be more important than capital outflows in affecting the dynamics of banks' funding structures.

Apart from the external financial liberalization measure, we also merge different macroeconomic variables to our bank-level data. Following Dinger and von Hagen (2009) and Gropp and Heider (2010), we expand our data set by real PPP adjusted per capita GDP (*PERCAPITAGDP*), the percent change in the consumer price index to control for the high inflation rates in many Latin American countries (*INFLATION*) and the real GDP growth rate (*GROWTH*). We further include the VIX as an additional covariate because it has been shown to be a good proxy for fluctuations in global risk aversion that drive the international supply of capital (e.g., te Kaat, forthcoming). Our macroeconomic data set also includes the unemployment rate, stock market volatility, the rule of law, the regulatory reserve and capital requirements and sovereign debt. Yet, as these variables turned out to be statistically insignificant in most of the regressions, we exclude them from the set of macro controls in the regression specifications reported in the remainder of this paper.[11] Table A.1 (Appendix) provides further specifics of the data.

3.3. Summary Statistics

Table 3 summarizes the main descriptive statistics of the bank-level and macroeconomic variables in our model. The mean bank has a capital-to-asset ratio of 18%, a deposit share of 52% and an interbank ratio of about 11%. These numbers show that, compared with advanced economies, banks in emerging markets fund a higher proportion of their balance sheet

[11] The insignificance of capital requirements is consistent with earlier research of Gropp and Heider (2010), who show that capital regulation only has a second order importance in determining banks' capital structures.

Table 3
Summary Statistics.

	Obs.	Unit	10th	Median	Mean	90th	SD
CAPITAL	8278	%	6.64	11.98	17.57	34.51	16.49
DEPOSITS	7637	%	8.94	58.68	52.16	81.64	26.04
INTERBANK	6676	%	0	4.75	11.17	31.73	15.49
OTHER SHORT-TERM DEBT	6238	%	0	4.59	11.40	33.58	15.85
BONDS	7106	%	0	3.61	10.25	29.50	16.00
NON-INTEREST FUNDS	8258	%	0.57	3.33	6.08	13.51	9.18
SIZE	8278	ln(x)	3.49	6.12	6.18	8.94	2.10
RISK	8278	%	−13.85	−0.39	5.10	27.58	30.99
NONINTERESTINCOME	8199	%	2.08	25.88	30.34	66.75	37.06
LIBERALIZATION	8278	-	0.2	0.80	0.67	1.00	0.31
PERCAPITAGDP	8278	-	5.35	11.39	11.27	17.16	4.29
INFLATION	8275	%	2.27	5.79	8.51	16.21	9.65
GROWTH	8278	%	−0.61	3.92	3.72	8.22	3.65
VIX	8278	%	12.81	22.55	21.64	31.48	5.99

The first six variables (the dependent variables employed in our analysis) are the bank-level shares of capital, retail deposits, interbank deposits, other short-term debt, bonds and non-interest funds in total assets. The bank controls added are the logarithm of total assets, impaired loans less reserves for impaired loans in equity and non-interest income over gross revenues. The macro covariates are the Schindler capital account inflows index, as well as per capita GDP, the inflation rate, real GDP growth and the VIX.

with equity and customer deposits, while interbank funding has a lower importance.[12] In addition, other short-term debt has an average share of 11%, bonds of 10%, and non-interest funding of 6%.

Turning to the other bank-level variables, Table 3 shows that the arithmetic mean of the variable *RISK* (defined as impaired loans less loan loss reserves over total equity) is equal to 5.1%, implying that, for the average Latin American bank, impaired loans exceed reserves for loan losses. There are also several banks with significant amounts of impaired loans, outstripping 27% of their equity (90th percentile). Further, the low mean ratio of non-interest income (30.34%) points to the fact that banks' business models in Latin America are focused on financial intermediation, so that interest income is the main source of revenue. The share of non-interest income over gross revenues in advanced economies is significantly higher and equal to 40% (e.g., DeYoung and Rice, 2004).

Our main regressor, capital account openness, has a mean value of 0.67. Thus, the average bank operates in a country which is externally relatively open. Yet, as pointed out before, the cross-country variation in this variable is far-reaching, including countries that are fully shielded from foreign capital (*LIBERALIZATION* = 0) and countries which are fully open (*LIBERALIZATION* = 1).

The values for per capita GDP in our sample vary substantially with a 10th percentile of 5,350 USD and a 90th percentile of 17,160 USD. The average inflation rate equals 8.51%. Non-trivial inflation stresses the great importance of controlling for changes in price levels, as they are likely to affect our estimates. Finally, the average real GDP growth rate is equal to almost 4% and the VIX takes a mean value of 21.64%.

4. Bank Funding Dynamics

4.1. Econometric Specification

We examine the relationship between changes in capital account regulations and banks' funding structures using the following model:

$$FUNDING_{ijt} = \alpha_i + \gamma * FUNDING_{i,j,t-1} + \beta * LIBERALIZATION_{jt} + \theta * X_{ijt} + \epsilon_{ijt} \qquad (1)$$

The dependent variables in Eq. (1) are the shares of capital, retail deposits, interbank funding, other short-term debt, bonds or non-interest liabilities over total assets of bank i in country j at time t. Most of the tables presented throughout the paper only show the results for the first three variables both because they are most critical for banks in Latin America (see Table 3) and because we harmonize the samples across the different dependent variables to make the effects of capital account openness more comparable and harmonizing across all of the six dependent variables would have reduced the number of observations to only 2690, thus lowering the precision of our estimates. Note, however, that Table A.2 contains the results for all six variables, using a non-harmonized sample. While the main results for capital, retail deposits and interbank funding are largely unaffected, capital account openness does not have a significant relation with other short-term debt, bonds or non-interest funds, justifying their exclusion from the main part of the paper. As our outcome variables exhibit non-trivial autocorrelation, it seems important to include the lagged dependent variables on the right hand side of equation (1) to help capture the time-series dynamics of

[12] In the euro area, for instance, the average share of customer deposits is equal to 30–40%, wholesale funding has a share of 20–30% and capital ratios are equal to about 6–8% (ECB (2016)).

banks' funding structures.[13] The coefficient α_i is an individual bank intercept and the vector X includes the bank-level and macroeconomic controls listed in Table 3. The main coefficient of interest in the following analysis is β, which measures the short-run impact of external financial liberalization on banks' funding ratios. The long-run effects are given by $\frac{\beta}{1-\gamma}$.[14]

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 (Nickell, 1981). To overcome these issues, we estimate the equation with the Blundell-Bond system GMM estimator (Blundell and Bond, 1998),[15] 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).

We instrument the regressors with five lags (lag 2-lag 6) of their levels and first differences, respectively.[16] Restricting the number of instruments is important because they increase quadratically in T and, therefore, can become very large, overfitting endogenous variables (Roodman, 2009b). The standard errors are corrected by the procedure proposed by Windmeijer (2005), which has been shown to address the potential downward bias of the two-step estimates of the system GMM standard errors that arises when using a large number of instruments in a regression. Its application makes our t-statistics more conservative. Finally, the regressions are weighted by banks' total assets. This is important in order to adjust our estimates for the oversampling of small banks, which are less of a concern from a financial stability/systemic risk perspective.

4.2. Baseline Results

As is apparent from Table 4, capital account openness is associated with reductions in banks' capital ratios and higher ratios of interbank funding. Retail deposits, in contrast, are not affected significantly by external financial liberalization. The same is true for other short-term debt, bonds and non-interest liabilities, as can be seen from Table A.2. In economic terms, an increase in the external financial liberalization index by one standard deviation (about 0.31 in our sample) reduces the capital-to-asset ratios on impact by 0.38 pp. The long-run effect is equal to 1.6 pp, as can be gauged by dividing the coefficient of *LIBERALIZATION* by (1-autoregressive coefficient). This is an economically significant effect since earlier research finds even smaller reductions in banks' equity ratios to increase the probability of bank distress significantly. For instance, the ECB (2015) finds that a one-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). Turning to the economic significance of interbank borrowing, Table 4 indicates that a one-sd increase in external financial liberalization raises banks' interbank deposits by 0.43 pp in the short-run; the long-run effect is equal to 2.8 pp.[17]

Banks' funding structures are also affected significantly by the set of bank-level controls. In particular, larger banks have lower equity ratios and less retail deposits. Risky banks and banks with lower non-interest income are also characterized by lower equity ratios. These results are consistent with earlier findings by Gropp and Heider (2010) or George (2015), among others. From the macroeconomic covariates, especially inflation rate differences affect banks' funding structures: a high inflation rate tends to raise the shares of retail and interbank deposits (which are typically of shorter maturities). Higher global uncertainty (higher VIX) is also associated with lower capital-to-asset ratios. Overall, in line with Gropp and Heider (2010), we find most other macroeconomic factors to be insignificantly associated with changes in bank funding structures.

In columns (1)-(3), the lagged dependent variables are highly statistically significant with a coefficient between 0.77 (for equity ratios) and 0.93 (for retail deposits). Therefore, retail deposits are more sticky (have higher autocorrelation) than other types of funding. These estimates further imply an adjustment speed (1-autoregressive coefficient) of 7%-23%. An adjustment speed of about 25% for banks' capital-to-asset ratios is broadly consistent with that obtained by Faulkender et al. (2012) and suggests that bank capital ratios adjust quickly.

4.3. Controlling for the On- vs. Off-Shore Interest Spread

We have previously shown that capital account openness leads to more interbank funding and less equity. This result is consistent with the notion that, in the wake of external financial integration, short-term debt flows are the dominant form of

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

[14] See, for instance, Koyck (1954) and Abbassi and Linzert (2012).

[15] We rely on the xtabond2 command in Stata (Roodman, 2009a) to estimate these regressions. For the interbank regressions, some specifications do reject the null of no second-order autocorrelation of the error terms. In these cases, we included a second lag of the dependent variable, which prevented the errors to be autocorrelated of order two, and the main results were largely unaffected. In order to be consistent with all other regressions in the paper, we do not report the attendant results, but they are available upon request.

[16] The results are robust to other lag specifications.

[17] Recalling the caveats on the exogeneity of capital controls discussed in Section 2, it can be easily seen that if capital account liberalization is partly endogenous to actual funding, this would bias the estimated coefficient downwards, not upwards. This is because bankers would, if anything, lobby to increase access to interbank borrowing when interbank ratios are low, thus entailing a negative covariance between liberalization and the funding variable. The same reasoning, in the opposite direction, holds for the coefficient on liberalization in the equity-to-asset regression, once again entailing a potentially stronger effect once endogeneity is fully removed. In Section 7, we instrument the liberalization variable to show that the effects become even slightly stronger.

Table 4
Baseline Results.

	(1) CAPITAL	(2) DEPOSITS	(3) INTERBANK
CAPITAL (t-1)	0.769***		
	(21.53)		
DEPOSITS (t-1)		0.926***	
		(55.92)	
INTERBANK (t-1)			0.847***
			(26.07)
LIBERALIZATION	−1.221***	0.123	1.398*
	(-3.41)	(0.14)	(1.86)
SIZE	−0.277***	−1.028***	0.151
	(-2.83)	(-3.47)	(0.83)
RISK	−0.006	0.008	−0.007
	(-1.61)	(0.42)	(-0.79)
NONINTERESTINCOME	0.010***	0.047***	−0.009
	(3.53)	(4.02)	(-1.26)
PER CAPITA GDP	−0.028	−0.006	−0.169***
	(-0.99)	(-0.07)	(-3.38)
INFLATION	−0.016	0.170***	0.022*
	(-1.55)	(4.84)	(1.71)
GROWTH	−0.041	0.032	0.048
	(-0.74)	(0.30)	(0.58)
VIX	−0.047***	0.014	−0.026
	(-3.12)	(0.21)	(-0.48)
Obs	4206	4206	4206
p (Hansen statistic)	0.996	0.995	0.998

The regressions are based on annual bank-level data over the period 1995–2013. The dependent variables are the shares of capital, retail deposits and interbank funding over total assets. The main regressor is the degree of capital account openness, proxied by the capital inflow index of Fernández et al. (2016). We further add several bank-level (the logarithm of total assets, the ratio of impaired loans and non-interest income in gross revenue) and macro (per capita GDP, inflation, GDP growth and the VIX) controls. The regressions are weighted by banks' total assets and estimated with the Blundell-Bond estimator, using five lags of the variables as instruments. We correct the standard errors by the procedure of Windmeijer (2005). The t-statistics are shown in parentheses and p (Hansen statistic) provides the p values for the Hansen test of overidentification restrictions.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

cross-border capital flows to emerging economies (e.g., Henry, 2007; Kose et al., 2009). Foreign investors, however, should provide disproportionately more short-term funding (i.e., interbank loans) to banks in emerging market regions the lower is the world interest rate. In this sub-section, we therefore expand the baseline analysis by testing whether external financial openness affects the funding structures of banks disproportionately more during periods of high real domestic money market interest rate spreads relative to the world's main financial center—the US. Since money market rates are mainly driven by the stance of monetary policy, the following analysis also allows us to analyze the interaction of capital account liberalization and monetary policy. To this end, the following analysis splits the sample into episodes in which the real domestic money market rate relative to the US is in the upper half of the in-sample distribution and those in which it is in the lower half.[18]

The attendant results are in line with our hypothesis, indicating that *LIBERALIZATION* has economically and statistically more significant effects during episodes of high domestic money market spreads (columns (4)-(6)). Economically, a one-sd increase in *LIBERALIZATION* during these episodes raises the interbank funding ratios by 0.63 pp in the short-run and by 4.6 pp in the long-run. The reduction in banks' equity ratios is equal to 0.56 pp on impact (2.3 pp in the long-run). These effects are 40%-60% larger than our baseline estimates. For interbank funding and equity ratios, a test of the equality of the estimates, following the z-test statistic proposed by Clogg et al. (1995), further rejects the null that both are equal to each other at the 5% level.

In a nutshell, we document that the effects of external financial liberalization on banks' funding structures are influenced by the stance of monetary policy at home and abroad. When money market rates in international financial centers relative to emerging market economies are low, capital account liberalizations in the latter are associated with a disproportionate decrease in equity and higher interbank funding. These results therefore add to the findings of recent work on the cross-border spill-overs of US monetary policy (Cetorelli and Goldberg, 2011, 2012; Bruno and Shin, 2015a; Ioannidou et al., 2015; Baskaya et al., 2017; Cerutti et al., 2017; Buch et al., 2019; Miranda-Agrippino and Rey, 2020), highlighting also a significant link between monetary policy in the US and the liability composition of banks in peripheral economies. In unreported regressions, we also examine the impact of capital account openness on bank profitability, proxied by their returns on assets. To the extent that banks are substituting cheaper foreign interbank lending for expensive funding from

[18] The data on money market rates come from the IMF's International Financial Statistics.

the local market, we should see that bank profitability increases. In fact, the coefficient estimate of a regression of returns on assets on capital account openness, after controlling for the set of macroeconomic and bank-level controls, is positive and statistically significant. This result is available upon request.

5. Are the Results Driven by Informationally Less Opaque Banks?

Following capital account liberalization, foreign investors mainly provide short-term debt funding, rather than equity, to borrowers in emerging markets (e.g., Henry, 2007; Kose et al., 2009). This result is attributed to asymmetric information between both parties. Due to such information asymmetries, the extant literature on the capital structures of non-financial corporates further shows that international/distant lenders prefer borrowers with rich information available to outside stakeholders (e.g., Lucey and Zhang, 2011). Next, we test whether this evidence on firms also applies to banks, i.e., whether capital account openness also benefits informationally less opaque banks disproportionately more. As these tests explore the differences across banks based on an interaction between a country (capital account openness) and a bank characteristic, the corresponding estimates are less sensitive to the underlying rationale for external financial liberalization, thus improving identification. For instance, even if unobservable macroeconomic variables correlate with *LIBERALIZATION*, inter-bank differences in the sensitivity with respect to external financial liberalization should be less affected.

As many empirical studies use size as a proxy for information availability, our first test explores the nexus between external financial integration and funding ratios conditional on bank size. If international investors tend to prefer lending to informationally less opaque banks, we should find a stronger effect of *LIBERALIZATION* on the funding structures of large banks. For the identification of this hypothesis, we enable *LIBERALIZATION* to interact with banks' logarithm of total assets, lagged by one year to minimize endogeneity concerns. As our regressions are already weighted by banks' total assets, this test basically examines whether, within the weighted sample of large banks, the **largest** financial institutions are affected most significantly by capital account liberalization. Attendant results, shown in columns (1)-(3) of Table 6, are broadly consistent with the aforementioned evidence on firms. Specifically, especially the shares of equity of the largest banks are affected by capital account liberalization. Whereas a one-sd increase in *LIBERALIZATION* even has a slightly positive impact on the equity shares of the median bank, the same effect for the largest banks at the 99th percentile of the distribution of the logarithm of total assets is equal to -0.63 pp. It also becomes apparent that for small banks, the impact of capital account openness on interbank funding ratios is insignificant, as can be seen from the statistically insignificant coefficient on *LIBERALIZATION*. However, although the interaction coefficient between *LIBERALIZATION* and *TOTAL ASSETS* is insignificant from a statistical point of view, it becomes obvious that the economic impact of capital account openness increases with bank size. In addition, once bank size reaches the 87th percentile of its distribution, the marginal impact of *LIBERALIZATION* on interbank ratios turns statistically significant. Thus, capital account liberalization and the improved access to foreign funding mainly affects the largest banks.

In the next set of tests, we examine whether the effects of *LIBERALIZATION* are amplified in foreign-owned banks, assuming that foreign ownership reduces the informational frictions between global investors and banks. For this analysis, we define foreign-owned banks as banks whose equity is to at least 50% owned by a foreign institution, using the ownership data provided in Claessens and van Horen (2014) and lagging the resulting foreign ownership dummy by one year.[19] Although columns (4) and (6) suggest that *LIBERALIZATION* has a stronger impact on the shares of equity and interbank deposits of foreign-owned banks, the corresponding interaction coefficients are not statistically significant at conventional significance levels. Thus, bank size in our empirical setting seems to be a better proxy for information availability than foreign ownership.

Previous regressions suggest that foreign investors overproportionally take positions in large Latin American banks, which are arguably subject to a lower degree of asymmetric information. Following this evidence, we finally allow the external financial liberalization index to interact with the one-year lag of the ratio of impaired loans less reserved for impaired loans relative to equity, a frequently used measure for the opaqueness of bank balance sheets.[20] We hypothesize that a more opaque balance sheet also increases the information asymmetries between domestic banks and international investors, thus reducing the effects of *LIBERALIZATION* on banks' funding structures. Columns (7)-(9) support this hypothesis: whereas the short-run effect of a one-sd increase in *LIBERALIZATION* on the interbank ratios of banks at the 10th percentile of the distribution of asset risk is equal to 0.81 pp, its effect on banks with impaired loans at the 90th percentile is only equal to 0.30 pp. The long-run difference in interbank ratios between both types of banks is even more pronounced (6.0 pp vs. 2.2 pp). We obtain a similar result for equity ratios, but the corresponding interaction coefficient in column (7) is not statistically significant. Overall, the results presented in this section thus indicate that the effects of external financial integration are amplified in banks with a lower degree of asymmetric information.

6. How Does Capital Account Liberalization Affect Small Banks?

Table 7 depicts the size distribution of banks in our data set. It shows that 90% of (smaller) banks have a combined asset share of less than 21%. In contrast, the largest 5% of banks in our sample have a combined asset share of 66.5%. As we

[19] We classify banks, which are not covered by Claessens and van Horen (2014), domestic.

[20] See Jungherr (2018). A higher share of impaired loans generally signals that the bank is prone to funding more opaque projects, 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).

Table 5
Controlling for the On- vs. Off-Shore Interest Spread.

	lower domestic interest spread			*higher domestic interest spread*		
	(1) CAPITAL	(2) DEPOSITS	(3) INTERBANK	(4) CAPITAL	(5) DEPOSITS	(6) INTERBANK
CAPITAL (t-1)	0.807***			0.760***		
	(22.26)			(17.49)		
DEPOSITS (t-1)		0.767***			0.934***	
		(18.78)			(57.36)	
INTERBANK (t-1)			0.775***			0.864***
			(14.65)			(24.00)
LIBERALIZATION	−0.610**	0.810	−0.165	−1.814***	−0.184	2.030**
	(-2.53)	(0.65)	(-0.29)	(-3.30)	(-0.15)	(2.11)
Bank Controls	Yes	Yes	Yes	Yes	Yes	Yes
Macro Controls	Yes	Yes	Yes	Yes	Yes	Yes
Obs	1765	1765	1765	1705	1705	1705

These regressions are based on annual bank-level data over the period 1995–2013. The dependent variables are the shares of equity, retail deposits and interbank deposits in total assets. The main regressor is the degree of capital account openness, measured by the capital inflow index of Fernández et al. (2016). We also add several bank-level (the logarithm of total assets, the ratios of impaired loans and non-interest income over gross revenue) and macro (per capita GDP, inflation, GDP growth and the VIX) controls. In the first three columns, we restrict the sample to episodes with a low domestic real money market rate relative to the US. Columns (4)- (6) restrict the sample to higher interest rate episodes. All the regressions are weighted by banks' total assets and estimated via the Blundell-Bond estimator, using five lags of the variables as instruments. We correct the standard errors by the procedure proposed in Windmeijer (2005). The t-statistics are shown in parentheses.
$* \; p < 0.10$, $** \; p < 0.05$, $*** \; p < 0.01$

Table 6
Are the Results Driven by Informationally Less Opaque Banks?

	(1) CAPITAL	(2) DEPOSITS	(3) INTERBANK	(4) CAPITAL	(5) DEPOSITS	(6) INTERBANK	(7) CAPITAL	(8) DEPOSITS	(9) INTERBANK
CAPITAL (t-1)	0.785***			0.785***			0.784***		
	(25.38)			(26.90)			(23.37)		
DEPOSITS (t-1)		0.945***			0.938***			0.931***	
		(69.60)			(60.88)			(63.74)	
INTERBANK (t-1)			0.849***			0.860***			0.866***
			(27.25)			(30.20)			(37.60)
LIBERALIZATION	3.094*	0.521	−3.174	−0.983**	0.870	1.047	−1.289***	−0.131	2.038**
	(1.95)	(0.14)	(-0.88)	(-2.15)	(0.80)	(1.06)	(-3.14)	(-0.16)	(2.44)
LIBERALIZATION × SIZE (t-1)	−0.461***	−0.096	0.465						
	(-2.61)	(-0.24)	(1.09)						
LIBERALIZATION × FOREIGN (t-1)				−1.235	−2.579	1.397			
				(-1.29)	(-0.81)	(0.66)			
LIBERALIZATION × RISK (t-1)							−0.013	0.016	−0.041**
							(-0.76)	(0.36)	(-2.03)
SIZE (t-1)	1.763***	6.348***	−2.454*						
	(2.81)	(3.16)	(-1.75)						
FOREIGN (t-1)				0.470	0.729	−0.593			
				(1.00)	(0.57)	(-0.72)			
RISK (t-1)							0.010	−0.015	0.005
							(0.66)	(-0.32)	(0.37)
Bank Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Macro Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Obs	4206	4206	4206	4206	4206	4206	3947	3947	3947

These regressions are based on annual bank-level data over the period 1995–2013. The dependent variables are the shares of equity, retail deposits and interbank deposits in total assets. The main regressor is the degree of capital account openness, proxied by the capital inflow index of Fernández et al. (2016), interacted with banks' logarithm of total assets, a foreign ownership dummy and the impaired loans ratios, respectively, all of which lagged by one year. We also add bank (the log of total assets, the ratio of impaired loans, non-interest income over gross revenue) and macro (per capita GDP, inflation, GDP growth, the VIX) controls. The regressions are weighted by banks' total assets and estimated with the Blundell-Bond estimator, using five lags of the variables as instruments. We correct the standard errors by the procedure proposed in Windmeijer (2005). The t-statistics are shown in parentheses.
$* \; p < 0.10$, $** \; p < 0.05$, $*** \; p < 0.01$

weighted all of the previous regressions by banks' total assets, we identified—to a great extent—the implications of capital account liberalization for the largest banks.

Table 7
The Distribution of Banks by Total Assets.

Size Class	Bank-Year Observations	Asset Share (in %)
< 50%	4139	1.8
50%-90%	3311	19.1
90%-95%	414	12.6
95%-99%	331	30.9
>99%	83	35.6

This table presents the number of bank-year observations for different bank size classes and the corresponding share of assets held by the particular size class (e.g., the first row shows the number of observations of the smallest 50 percent of banks in our sample, as well as their total assets relative to aggregate total assets of the whole banking system).

Table 8
The Effects of Capital Account Liberalization on Small vs. Large Banks.

	(1) CAPITAL	(2) DEPOSITS	(3) INTERBANK	(4) DEPOSIT RATE, ALL BANKS
CAPITAL (t-1)	0.779***			
	(18.93)			
DEPOSITS (t-1)		0.873***		
		(36.64)		
INTERBANK (t-1)			0.749***	
			(26.75)	
DEPOSIT INTEREST RATE (t-1)				0.805***
				(32.80)
LIBERALIZATION	−0.522	2.264**	−2.144***	1.425
	(-0.94)	(2.10)	(-2.83)	(0.59)
LIBERALIZATION × BIG (t-1)				−5.877*
				(-1.87)
BIG (t-1)				0.550
				(0.31)
Bank Controls	Yes	Yes	Yes	Yes
Macro Controls	Yes	Yes	Yes	Yes
Obs	2947	2947	2947	3855

Columns (1)-(4) are based on annual bank-level data over the period 1995–2013. The dependent variables are the shares of equity, retail and interbank deposits in toal assets and the deposit interest rate. The key regressor is capital account openness measured by the capital inflow index of Fernández et al. (2016). We add several bank-level (the logarithm of total assets, the ratios of impaired loans and non-interest income in gross revenues) and macro (per capita GDP, inflation, GDP growth and the VIX) covariates. The regressions are estimated via the Blundell-Bond estimator, using five lags of the variables as instruments. We correct the standard errors by the procedure proposed in Windmeijer (2005). The t-statistics are shown in parentheses. The regression in column (4) is weighted by banks' total assets and we also interact LIBERALIZATION with a dummy equal to one for banks larger than the 95th percentile of the size distribution. Columns (1)-(3) drop the largest 25% of banks in terms of total assets from the sample.
$* \, p < 0.10$, $** \, p < 0.05$, $*** \, p < 0.01$

In this section, we document whether and through which channels external financial openness also affects small banks' funding dynamics, which is important because small banks are typically the main provider of credit to small/more opaque non-financial firms (Berger and Udell, 2002). To this end, we refrain from weighting the observations by total assets in the regressions presented in Table 8 and we drop the largest 25% of banks in terms of total assets.[21] Unlike our baseline analysis, capital account openness leads to higher shares of retail deposits (column (2)) and lower interbank ratios (column (3)) for small banks. These effects are economically meaningful: in the long-run, the shares of retail funding increase by 5.5 pp and banks' interbank ratios decrease by 2.6 pp in the wake of a one-sd increase in *LIBERALIZATION*.

The rise in interbank borrowing for the largest banks and higher retail deposits for small banks, in turn, raise the question on the transmission of global liquidity to the different types of banks. We conjecture that, in response to the lower relative cost of foreign interbank borrowing, large banks lower their deposit interest rate relative to that of smaller banks, inducing deposit flows to the latter and, thus, making the latter more dependent on retail funding.[22] To verify this hypothesis, we continue regressing the deposit interest rate, defined as interest expenses on retail deposits over total retail deposits, on the interaction between capital account openness and a lagged dummy that is equal to one for banks in the top 95% of the country-year-specific in-sample distribution of total assets, zero otherwise.[23] Column (4) shows that external financial liberalization induces

[21] Dropping a smaller fraction of banks leaves the results qualitatively unchanged, but reduces the statistical significance of the coefficients.

[22] This hypothesis is broadly in line with the extant literature on the link between banks' market power and deposit rates, which shows that smaller (single-market) banks depend disproportionately more on customer deposits and, as a result, attract retail funding by paying a higher relative deposit interest rate (e.g., Barros, 1999; Hannan and Prager, 2006; Park and Pennacchi, 2009).

[23] We calculate this threshold using a country-specific distribution because competition for retail deposits is likely to happen only within a country. The results are robust to employing alternative thresholds, such as the 98th percentile, to differentiate between small and large banks. Due to some extreme outliers in the deposit interest rate, we winsorize this variable at the 95% level in order to ensure that our results are not driven by unrepresentative outliers.

large banks to lower their deposit interest rate relative to that of smaller banks, indicated by a negative coefficient on the interaction between *LIBERALIZATION* and the large bank dummy, which is statistically significant at the 10% level. In economic terms, whereas the deposit rate for small banks is unaffected by capital account openness, indicated by an insignificant coefficient of *LIBERALIZATION*, a one-sd increase in external financial openness reduces the retail deposit rate of large banks by 1.4 pp, which is non-trivial given a median deposit interest rate of 8.4% in our sample. Overall, we therefore establish that small banks benefit indirectly from capital account openness via an increased access to retail deposits.

7. Robustness Checks

In this section, we present several additional results, including those of various robustness checks. In a first set of regressions, we do not harmonize the sample so that the regressions for each dependent variable have a different number of observations. We also present the results for regressions of other short-term debt, bonds and non-interest liabilities on capital account openness. These variables have been excluded from Table 4 because harmonizing across all of the six dependent variables would have reduced the number of observations to only 2690 and thus lowered the precision of our estimates. As can be seen from Table A.2, our baseline estimates are largely unaffected when we refrain from homogenizing the sample at all. If anything, the statistical significance increases relative to Table 4. In addition, Table A.2 shows that other short-term debt, bonds and non-interest liabilities are not affected significantly by *LIBERALIZATION*.

We continue estimating Eq. (1) via ordinary least squares. As is apparent from Table A.3, capital account openness is still associated with significantly higher interbank funding ratios. Further, *LIBERALIZATION* also reduces banks' capital ratios. As in our baseline analysis, there is no significant link between capital account openness and the shares of retail deposits. Thus, our main results are robust to OLS estimation.

Next, we estimate our model with the Blundell-Bond estimator, but include a government's partisanship indicator and an IMF program dummy as exogenous instruments in the estimation. Both variables are exogenous to external financial liberalization and, additionally, significant drivers of the latter (see the discussion in Section 2). We are thus able to improve identification. Table A.4 corroborates our baseline estimates: a one-sd increase in external financial liberalization in the short-term increases banks' interbank ratios by 0.46 pp and reduces banks' capital ratios by 0.40 pp. Retail deposits, in contrast, are not affected by external financial liberalization at conventional significance levels.

Finally, we adjust the time coverage of our sample by estimating equation (1) over different sub-samples. First, we differentiate between the pre-2008 and post-2008 period to examine whether the relation between capital account openness and banks' funding ratios differs in the pre- and post-global financial crisis episode. Second, we drop the years before 1999. Although we lose some variation in the international financial liberalization measure, this adjustment might be important because the Bankscope database has a higher coverage for the period 1999–2013 (see Table 1).[24] The attendant results in Table A.5 show that the effects of external financial openness are stronger post-2008 than pre-2008, but the coefficients are also estimated less precisely. In addition, columns (7)-(9) indicate that *LIBERALIZATION* is still associated with lower equity ratios and more interbank borrowing once we exclude the years with worse Bankscope coverage.

8. Concluding Remarks

To the best of our knowledge, this is the first paper that relates changes in capital account controls to the various liability components of banks employing bank-level data and a sizable panel of emerging market economies. Recent research on the effects of international financial integration on the banking systems of emerging economies has focused on the asset side of banks, falling short of a detailed analysis of the effects of capital control changes on the composition of bank liabilities. This paper shows that this neglect is unwarranted, since relaxations of capital controls are associated with a substitution of interbank funding for equity among large banks—an effect that is likely to dominate at the macro level due to the size concentration of the domestic banking systems in many emerging market economies. We also show that this substitution is more significant among informationally less opaque banks. Further, such effects are stronger during low global interest rate episodes.

Many emerging market countries still impose relatively stringent capital controls. The findings of this paper thus highlight a policy-relevant and hitherto overlooked mechanism through which further relaxations of capital controls can increase the propensity for financial instability: all else constant, large banks tend to increase their reliance on short-term interbank funding, boosted by external liquidity, and increase their leverage; meanwhile, the funding of smaller banks becomes more dependent on interest-sensitive deposits, which are prone to flight-to-safety once systemic shocks hit. Thus, to the extent that capital account openness makes the interbank market more vulnerable to sudden stops in capital inflows, increasing the funding risk of large banks directly and that of small banks indirectly, it thereby generates externalities that make the consolidated financial system more vulnerable to rollover risk. These findings should not be interpreted as a rejection of the many benefits from international financial integration, but they do suggest that macro-prudential regulations have a role to play as countries open up their capital accounts.

[24] 1999 is the first year where the number of banks exceeds 350.

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Appendix A

Tables A.1,A.2,A.3,A.4,A.5.

Table A.1
Description of the Variables.

Variable	Description	Unit	Source
CAPITAL	equity/ total assets	%	Bankscope, own calculations
DEPOSITS	customer deposits/ total assets	%	Bankscope, own calculations
INTERBANK	interbank liabilities/ total assets	%	Bankscope, own calculations
OTHER SHORT-TERM DEBT	other short-term debt/ total assets	%	Bankscope, own calculations
BONDS	traded liabilities/ total assets	%	Bankscope, own calculations
NON-INTEREST FUNDS	non interest-bearing liabilities/ total assets	%	Bankscope, own calculations
SIZE	ln (total assets)	ln (million x)	Bankscope, own calculations
RISK	(impaired loans - reserves for impaired loans)/ equity	%	Bankscope, own calculations
NONINTERESTINCOME	non-interest income/ gross revenues	%	Bankscope, own calculations
FOREIGN	=1 if bank equity is to at least 50% owned by a foreign institution	0/1	Claessens and van Horen (2014), own calculations
TOTAL ASSETS	total assets	billion x	Bankscope, own calculations
LIBERALIZATION	(1 - Schindler inflow restrictions index)	-	Fernández et al. (2016), own calculations
PERCAPITAGDP	PPP adjusted per capita GDP	x/1000	WEO, own calculations[a]
INFLATION	The relative change in the CPI index	%	WEO, own calculations
GROWTH	The real GDP growth rate	%	WEO
VIX	The CBOE Volatility Index	%	Chicago Board Options Exchange

[a] World Economic Outlook Database, IMF.

Table A.2
Baseline Results Without Sample Harmonization.

	(1) CAPITAL	(2) DEPOSITS	(3) INTERBANK	(4) OTHER SHORT-TERM DEBT	(5) BONDS	(6) NON-INTEREST FUNDS
CAPITAL (t-1)	0.709***					
	(29.44)					
DEPOSITS (t-1)		0.935***				
		(71.97)				
INTERBANK (t-1)			0.861***			
			(33.15)			
OTHER SHORT-TERM DEBT (t-1)				0.776***		
				(18.30)		
BONDS (t-1)					0.924***	
					(39.26)	
NON-INTEREST FUNDS (t-1)						0.757***
						(15.19)
LIBERALIZATION	−1.117***	0.216	1.296**	0.213	−0.109	1.038
	(-3.45)	(0.29)	(2.11)	(0.25)	(-0.009)	(1.48)
Obs	6877	6370	5383	4918	5694	6854

The regressions are based on annual bank- level data over the period 1995–2013. The dependent variables are the shares of capital, retail deposits, interbank funding, other short-term debt, bonds and non-interest liabilities over total assets. The main regressor is the degree of capital account openness, proxied by the capital inflow index of Fernández et al. (2016). We further add several bank-level (the logarithm of total assets, the ratio of impaired loans and non-interest income in gross revenue) and macro (per capita GDP, inflation, GDP growth and the VIX) controls. The regressions are weighted by banks' total assets and estimated with the Blundell-Bond estimator, using five lags of the variables as instruments. We correct the standard errors by the procedure of Windmeijer (2005). The t-statistics are shown in parentheses.
$* \; p < 0.10$, $** \; p < 0.05$, $*** \; p < 0.01$

Table A.3
OLS Results.

	(1) CAPITAL	(2) DEPOSITS	(3) INTERBANK
CAPITAL (t-1)	0.880***		
	(55.93)		
DEPOSITS (t-1)		0.942***	
		(84.04)	
INTERBANK (t-1)			0.887***
			(30.39)
LIBERALIZATION	−0.692*	0.477	1.169*
	(-1.68)	(0.48)	(1.94)
Bank Controls	Yes	Yes	Yes
Macro Controls	Yes	Yes	Yes
Obs	4206	4206	4206

These specifications are based on annual bank-level data for the period 1995–2013. The dependent variables are the shares of capital, retail deposits and interbank deposits over total assets. The main regressor is the degree of capital account openness, proxied by 1- inflow index of Fernández et al. (2016). We include bank-level (the logarithm of total assets, the fraction of impaired loans and non-interest income over gross revenue) and macro (per capita GDP, VIX inflation and GDP growth) controls. All the regressions are weighted by banks' total assets and estimated via OLS. The t-statistics are presented in parentheses employing heteroskedasticity robust standard errors.
$* \ p < 0.10$, $** \ p < 0.05$, $*** \ p < 0.01$

Table A.4
Instrumenting Capital Account Openness.

	(1) CAPITAL	(2) DEPOSITS	(3) INTERBANK
CAPITAL (t-1)	0.768***		
	(20.44)		
DEPOSITS (t-1)		0.922***	
		(54.42)	
INTERBANK (t-1)			0.848***
			(26.44)
LIBERALIZATION	−1.303***	−0.137	1.491*
	(-3.55)	(-0.14)	(1.81)
Bank Controls	Yes	Yes	Yes
Macro Controls	Yes	Yes	Yes
Obs	4206	4206	4206

These specifications are based on annual bank- level data for the period 1995–2013. The dependent variables are the shares of capital, retail deposits and interbank deposits over total assets. The main regressor is the degree of capital account openness, proxied by 1- inflow index of Fernández et al. (2016). We include bank-level (the logarithm of total assets, the fraction of impaired loans and non-interest income over gross revenue) and macro (per capita GDP, VIX inflation and GDP growth) controls. All the regressions are weighted by banks' total assets and estimated with the Blundell-Bond estimator; however, we include an IMF program dummy and a partisanship indicator as exogenous instruments. We correct the standard errors by the method of Windmeijer (2005). The t-statistics are shown in parentheses.
$* \ p < 0.10$, $** \ p < 0.05$, $*** \ p < 0.01$

Table A.5
Excluding Certain Time Periods.

	pre-2008			*post-2008*			*only observations post 1998*		
	(1) CAPITAL	(2) DEPOSITS	(3) INTERBANK	(4) CAPITAL	(5) DEPOSITS	(6) INTERBANK	(7) CAPITAL	(8) DEPOSITS	(9) INTERBANK
CAPITAL (t-1)	0.853***			0.692***			0.767***		
	(29.47)			(11.31)			(21.07)		
DEPOSITS (t-1)		0.929***			0.918***			0.925***	
		(42.30)			(39.62)			(52.70)	
INTERBANK (t-1)			0.936***			0.801***			0.844***
			(15.51)			(15.45)			(24.83)
LIBERALIZATION	−0.335	−1.263	1.777*	−2.412***	0.256	2.305	−1.225***	0.501	1.432*
	(-1.07)	(-0.72)	(1.75)	(-4.43)	(0.20)	(1.59)	(-3.24)	(0.56)	(1.82)
Bank Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Macro Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Obs	2300	2300	2300	1906	1906	1906	4019	4019	4019

These regressions are based on annual bank-level data over the period 1995–2013. The dependent variables are the shares of equity, retail deposits and interbank funding over total assets. The key regressor is the degree of capital account openness, proxied by 1- capital inflow index of Fernández et al. (2016). We further add several bank (the logarithm of total assets, the ratio of impaired loans and non-interest income to gross revenue) and macro (per capita GDP, inflation rate, GDP growth, VIX) controls. We restrict the sample to pre-2008 (columns (1)-(3)), post-2008 (columns (4)-(6)) and the period of 1999–2013 (columns (7)-(9)). The regressions are weighted by banks' total assets and estimated via the Blundell-Bond estimator with five lags of the variables as instruments. We correct the standard errors by the procedure proposed in Windmeijer (2005). The t-statistics are shown in parentheses.
$* \ p < 0.10$, $** \ p < 0.05$, $*** \ p < 0.01$

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