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Holes in the dike: the global savings glut, U.S. house prices and the long shadow of banking deregulation*

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Abstract

We show how capital inflows into and financial deregulation within the United States interacted in driving the recent boom and bust in U.S. housing prices. Interstate banking deregulation during the 1980s cast a long shadow: in states that opened their banking markets to out-of-state banks earlier, house prices were more sensitive to aggregate U.S. capital inflows during 1997-2012. Capital inflows relaxed the value-at-risk constraints of geographically diversified ("integrated") U.S. banks more than those of local banks. Therefore, integrated banks absorbed a larger share of capital inflows and expanded mortgage lending more. This drove up housing prices.

Keywords: house prices, global imbalances, interstate banking deregulation

JEL classification: G10, G21, G28, F20, F32, F40

1. Introduction

In this paper, we investigate empirically how the interaction between financial liberalization and capital inflows contributed to the rise (and subsequent decline) in U.S. residential housing prices. Our analysis makes use of ex ante differences in financial openness between U.S. federal states. During the late 1970s and 1980s, federal states lowered the legal barriers to access to their local banking markets (see Kroszner and Strahan (1999); Jayaratne and Strahan (1996)). We show that house prices in states that had lowered these barriers earlier were more sensitive to the impact of international capital flowing into the United States in the second half of the 1990s. Over the decade between 1995 and 2005 the U.S. was running a current account deficit of 3.5 percent of GDP on average. Our estimates imply that due to these capital inflows house prices at the peak of the housing boom in 2005 were more than 25 percent higher in a state that liberalized, say, in 1985 than in a state that liberalized six years later, in 1991.

Already prior to the global financial crisis of 2007/2008, some analysts saw global imbalances in capital flows—the high savings rates of emerging economies and the large current account deficits of the U.S. economy—as a major threat to global economic stability (Obstfeld and Rogoff (2009)). As early as 2005, Federal Reserve Chairman Ben Bernanke argued that a global glut of savings flowing into the U.S. was lowering long-term interest rates and thus contributing to a run-up in asset prices (Bernanke (2005)). However, when the financial crisis eventually struck in 2007, it emanated from what at first appeared as the least globalized part of the U.S. financial

system—the housing market and the market for residential mortgages. Many policymakers and academics therefore singled out the financial liberalization and deregulation of the two decades before 2007 as the main causes of the crisis. Only recently research has started to rigorously investigate the possibility that global imbalances in capital flows could themselves have contributed to the boom and bust in asset prices, notably in the price of housing (Aizenman and Jinjara (2009); Geta (2009); Favilukis et al. (2012); Justiniano et al. (2013); Punzi (2013); Ferrero (2015); Cesa-Bianchi et al. (2018)).

Our empirical contribution is to exploit the interaction between state-level histories of financial liberalization and aggregate U.S. capital inflows to identify the channels through which global imbalances affected housing prices. The intuition underlying our analysis is simple: when U.S. federal states deregulated during the 1980s, they effectively poked holes in the dikes that shielded their local banking markets from banking flows from outside the state. After liberalization, out-of-state banks started to enter these freshly liberalized local banking markets. States that liberalized earlier had a larger presence of banks operating in several states (referred to in what follows as “geographically diversified” or “integrated” banks) when global imbalances started to hit the U.S. from the mid-1990s onward as a consequence of the appearance of China and other emerging economies on the world economic stage (see Obstfeld and Rogoff (2009)). We show that the savings glut was a positive funding shock to the entire U.S. banking system. Geographically more diversified banks increased their leverage more, thus absorbing a larger share of the international demand for U.S. assets. States with a stronger presence of integrated banks—i.e. states that had poked bigger holes into their

dikes by liberalizing earlier during the 1980s—therefore were more exposed to the savings glut and expanded lending more. This drove up housing prices. We document this mechanism at different levels of aggregation: at the state level as well as based on granular bank- and county-level data.

We start by documenting our results on quarterly state-level data. During the period of global imbalances, that we date from 1997 to 2012, the correlation between aggregate capital inflows and subsequent house price growth is higher for states that opened their banking markets earlier during the liberalization wave of the 1980s. This result is robust to controlling for differential pre-sample trends in terms of a host of state characteristics that are likely to be correlated with financial liberalization and that could also explain inter-state differences in the sensitivity of house prices to capital inflows. It also holds when we control for a range of credit supply and demand factors, such as monetary policy, interest rate spreads and variation in house price expectations as well as for a second wave of bank branching liberalizations that rippled through the U.S. from the mid-1990s.¹

The general idea behind our approach is that aggregate inflows into the U.S. are driven by U.S.-wide factors and therefore largely exogenous with respect to house prices in individual states. Furthermore, the wave of cheap

1. Favilukis et al. (2017) model the role of domestic financial liberalization as a joint driver for capital inflows and house prices. Adelino et al. (2016); Cheng et al. (2014); Foote et al. (2012) explore the role of demand, in particular of house price expectations in the recent U.S. housing boom. Taylor (2007); Del Negro and Otrok (2007) document the role of monetary policy for house price dynamics and Jorda et al. (2015) present historical evidence on the link between interest rates and house prices.

capital in hunt for safe assets that hit the U.S. in the second half of the 1990s could not have been anticipated by state regulators in the 1980s, when most of the liberalization of the interstate banking regime took place. Thus, our primary measure of ex ante financial openness—the number of years that had passed since a state liberalized its interstate banking regime until 1995—is clearly pre-determined. Our state-level results therefore suggest that financial liberalization cast a long shadow on the way in which different states absorbed the savings glut. But a full causal interpretation of these results would require that the interaction between a state’s liberalization history and aggregate U.S. capital inflows is truly exogenous (after appropriate controls) to state- or county-level demand for housing and mortgages. In spite of all the controls discussed above, state level data are too aggregated to allow us to be fully confident that we have accounted for all state characteristics that could correlate with financial liberalization and that could also modulate the impact of aggregate capital inflows.

We therefore address remaining challenges to identification in the second part of the paper where we examine the impact of capital inflows on mortgage credit supply directly at the bank- and county-level. We obtain data from the Home Mortgage Disclosure Act (HMDA) data base which allows us to compute the growth rate of mortgage lending for each U.S. bank in each U.S. county. We then adapt the method recently proposed by Amiti and Weinstein (2018) for matched bank-firm data to our setting, with counties borrowing from different banks taking the place of firms. This new method allows us to disentangle bank-specific shocks to mortgage supply from county-specific shocks to mortgage demand. The method explicitly takes account of the

granular structure of banking markets in which the bulk of aggregate lending is concentrated among just a few big financial institutions (Landier et al. (2017); Bremus et al. (2018)). The shocks identified by the Amiti and Weinstein (2018) method are the only ones that aggregate correctly to match the growth of mortgage lending at the level of individual banks and counties.

To structure the empirical analysis at the bank-level, we provide a simple theoretical model based on Shin (2010) in which bank lending supply is determined by a value-at-risk (VaR) constraint. Due to geographical diversification, integrated banks' notional loan portfolio risk is lower than that of local banks. This is consistent with the widely-held view in the years before 2007 that the U.S.-wide housing market—in spite of some regional declines—would never decline in aggregate. In this framework, we then interpret the savings glut as a positive deposit supply shock to the U.S. banking system. Capital inflows drive down banks' refinancing rates which in turn relaxes their VaR constraints. Because integrated banks' perceived portfolio risk is lower, their leverage is higher and—importantly—also more sensitive to declines in refinancing rates than that of local banks. Thus, the model predicts that banks' elasticity of loan supply with respect to aggregate capital inflows increases with the geographical diversification of their mortgage portfolio.

We find strong support for this mechanism in the data. Our bank-county data set allows us to compute measures of the geographical diversification of banks' mortgage portfolios. Controlling for other economy-wide credit supply factors, bank size and for unobserved bank heterogeneity, we find that the elasticity of the bank credit supply shocks identified from the Amiti and

Weinstein (2018) procedure to aggregate capital inflows increases directly with banks' geographical diversification.

Having identified the causal impact of capital inflows on bank credit supply, we move on to construct county-level capital inflow shocks. We construct these as the weighted average of the capital inflow shocks to banks operating in a county. However, using a banks' de facto market share in a given county as weight in this aggregation could be problematic because these market shares could themselves be endogenous to differences in local house price expectations and other local-specific demand factors. To address this issue, we take guidance from our state-level results by exploiting a particular feature of the history of interstate banking deregulation in the United States: states that opened their markets for out-of-state banks generally did so on a mutually reciprocal basis, i.e. out-of-state banks were allowed to enter only if their home state allowed entry as well. This feature generated a wide variation in opportunities for entry among banks headquartered in different states. We exploit this variation to construct "as-if" market shares for banks. Specifically, we assume that banks' lending in a given county is a linear function of the time that the bank would have been allowed to enter markets in the state in which the county is located. If there are multiple banks active in the county and these banks have different home states, the implied market share of the bank becomes a complicated function of the liberalization history of the banks' home state, the target state it is operating in, and the liberalization histories of all states in which other banks operating in the county are headquartered. Our identifying assumption is that this imputed regulatory or de iure market share is orthogonal to local (county- or state-level) credit demand. We then

use these de iure weights to aggregate the bank-specific capital inflow shocks to the county level.

We show that this “granular” county-level capital inflow shock has considerable predictive power for county-level house prices after controlling for a host of potentially confounding factors. In particular, our results again hold up when we control for the gradual dismantling of interstate branching deregulation as well as for state-time specific effects more generally. Consistent with our theoretical framework, the granular capital inflow shock strongly predicts growth of mortgage lending by banks, while it is uncorrelated with the lending by non-banks such as e.g. independent mortgage companies. We therefore also use the granular county-level capital inflow shock as an instrument for lending in regressions of house prices on lending growth. Again, our results stand. In addition, we provide independent evidence based on another data set—banks’ call reports—showing that geographically more diversified banks saw lower refinancing rates as a result of the savings glut. This allowed them to increase their leverage and to also lower their lending rates. The fact that capital inflows lower the borrowing rates of geographically diversified banks is the ultimate indication that capital inflows acted as a liquidity supply shock to the U.S. banking system. We note that all our bank-level results hold while controlling for bank size and unobserved bank heterogeneity that could correlate with geographical diversification.

A final buttress for the causal interpretation of our results is provided by re-running our bank-level analysis on a placebo sample of non-bank, non-depository financial institutions—mortgage companies. As highlighted

by Demyanyk and Loutskina (2016), mortgage companies are exempt from the specific regulation that governs the risk-taking behavior of depository institutions. For this reason and because they primarily originate loans to sell them for securitization, the geographical diversification of mortgage companies should not matter for their lending supply response to capital inflow shocks. This is exactly what we find in the data.

Our paper is related to and builds on several important contributions that have recently explored the causal link between banking deregulation and house prices. Favara and Imbs (2015) document that state-level branching deregulation in the second half of the 1990s impacted house prices. Landier et al. (2017) show that liberalization of state-level banking markets increased the synchronization of housing prices between states. Mian et al. (2017) explore the channels through which state-level liberalization in the 1980s impacted credit supply. They find that deregulation predominantly affected consumer credit and show how this contributed to the housing boom and bust of the late 1980s and early 1990s. Our analysis complements these studies along several important dimensions. First, different from all three studies, we draw attention to the role of international capital flows in driving house prices. We also document that global imbalances were a major common factor in the lending decisions of geographically diversified banks and thus contributed significantly to the regional synchronization of house price developments. Secondly, we focus on a new channel of transmission by emphasizing how leverage targeting and better (notional) diversification of the asset side of their balance sheets allowed integrated banks to tap the global demand for U.S. safe assets. Third, we document a long shadow of banking deregulation by

showing that the impact of capital inflows on house prices in the second half of the 1990s was pre-determined by a state's liberalization history more than a decade earlier.

We also consider to what extent capital inflows may have contributed to financial instability by fueling excessive lending and house price increases in the run-up to the financial crisis. Starting with Mian and Sufi (2009), the literature has documented that the securitization of mortgage debt provided lenders with perverse incentives for lax screening and excessive lending. We find that banks securitized less in states that had a stronger exposure to capital inflow shocks. This suggests that a long history of openness may actually have helped improve the screening of credit quality in a state and may thus have reduced the abuse of securitization.

The idea that global demand for safe assets was a key driver behind global capital flows was first articulated theoretically by Caballero et al. (2008). Caballero and Krishnamurthy (2009) discuss a model in which the global demand for safe assets drives the prices of risky assets by allowing the domestic financial sector to increase leverage. Our results lend strong empirical support to a particular variant of this mechanism: integrated banks benefited from the global demand for safe assets to leverage up on risky assets (mortgages) because their geographical diversification allowed them to effectively turn their balance sheets into private-label safe assets backed by the entire U.S. housing market. Our findings, therefore, suggest that the intra-national liberalization of U.S. banking markets during the 1980s cast a long shadow: it helped lay the foundation for the ability of the U.S. financial system to provide more safe assets when the demand of emerging economies

for these assets surged roughly a decade later, triggered by the aftermath of the Asian crisis.

Borio and Disyatat (2011), Adrian and Shin (2010b) and Shin (2012) have pointed at the role of international gross banking flows in the build-up of leverage in the financial system in the years before the subprime crisis. Bruno and Shin (2015) emphasize the role of bank leverage for the international transmission of monetary policy. Our results show that banks' leverage was also key for the transmission of a general refinancing shock (the savings glut) across U.S. states: global imbalances in net flows drove down refinancing costs and the leverage of U.S. banks with lower notional risk (i.e. better regional diversification) reacted more to this cheaper funding than that of less diversified (local) banks. This mirrors the behavior of international banks that took advantage of cheap dollar funding to invest into a geographically diversified portfolio of (securitized) U.S. mortgages. We provide some results that suggest that net capital inflows indeed changed banks' incentives for geographical diversification—be it directly through an expansion of branch networks or indirectly through securitization and a concomitant increase in international cross-holdings of securitized assets.

The paper is structured as follows. Section two provides some historical background on U.S. state-level banking deregulation. Section three presents our state-level analysis. Section four introduces a simple theoretical framework that allows us to interpret our findings. Section five tests our theoretical framework on granular bank- and county-level data. Section six concludes.

2. State-level segmentation of U.S. banking markets: a brief history

Our analysis exploits the gradual dismantling of geographical restrictions on interstate bank expansion in the United States during the 1980s and early 1990s. These restrictions dated back to the 19th century, when states acquired the right to levy bank-licensing fees and generally prohibited out-of-state banks from operating in their territories. The McFadden Act of 1927 reaffirmed the authority of states over national banks' branching within their borders. However, at the same time it opened the possibility for geographical expansion through means of a "bank holding company" (BHC): a BHC can, in principle, operate banks in several states, as long as these banks remain separately capitalized legal entities.

The Douglas Amendment to the Bank Holding Company Act of 1956 gave states even stronger authority to prohibit out-of-state banks from acquiring banks outside the state where they were headquartered. Since all states implemented this prohibition, interstate banking in the U.S. was effectively barred from the mid-1950s until the late 1970s, when this regulation was gradually starting to be diluted. Beginning with Maine in 1978, state legislatures began to enact laws that allowed out-of-state BHCs to control banks in their state. Initially, such statutes authorized out-of-state acquisition only on a reciprocal basis with like-minded states or insisted that acquirers be headquartered in a neighboring state. Furthermore, federal legislators amended the Bank Holding Company Act in 1982 to allow failed banks to be acquired by any holding company, regardless of state laws. Over the following 13 years, states removed entry restrictions for bank holding companies by

unilaterally opening their state borders and allowing out-of-state banks to enter, or by signing reciprocal bilateral and multilateral agreements with other states to allow interstate banking. As the last state, Hawaii passed reciprocal entry laws in 1995.

It is important to note that all of these deregulations still did not allow full interstate bank branching but only the ownership and operation of local banks by out-of-state BHCs. Full branching was only implemented by provisions in the Riegle-Neal Interstate Banking and Branching Efficiency Act of 1994 which became effective in 1997. But states could opt out even from these interstate branching provisions—and many did, only gradually relaxing bank branching restrictions for out-of-state banks over the following decade. In all our regressions below we therefore control for the impact of interstate bank-branching deregulation during the 1990s in the context of our analysis. Our finding that the impact of capital inflows on house prices is stronger in states that liberalized earlier, is robust to controlling for this second wave of liberalizations.² In fact, our theory and some empirical results suggest that this second wave could partly be a response to the capital inflows that hit the U.S. during that time.

2. In many states intra-state branching was prohibited until the late 1980s. Thus, even banks from within the state were often confined to one county. It is important to distinguish the wave of interstate branching deregulations during the second half of the 1990s from this intra-state branching deregulation a decade earlier. At the level of individual states, the intra-state deregulation of the 1980s in most cases took place ahead of the interstate-liberalization discussed in the main text above and which is our focus here.

Since states deregulated in waves, or cohorts, rather than all at once, the staggered timing of interstate banking deregulation during the 1980s provides an ideal laboratory to explore empirically how these regulatory differences in openness to a bank entry affected the real economy. In our empirical analysis, we generally measure a state's banking openness as the number of years that have passed between the year in which the state allowed full interstate banking (through BHCs) and 1995. The important feature of our analysis is that the era of interstate liberalization in the 1980s precedes the era of global imbalances which started in the mid-1990s and reached its peak between 1997 and 2008, when in particular Asian economies started to accumulate international reserves on a previously unseen scale in the wake of the region's 1996/1997 financial crisis. Our measure of openness therefore is clearly predetermined with respect to the major wave of capital that hit the U.S. from the mid 1990s onward.

3. State-level analysis

3.1. *A first look at the data*

Our main measure of (net) capital inflows is the (negative) ratio of the current account to total equity of all commercial banks in the United States, i.e.

$$\text{CAPFLOW}_t = -\frac{CA_t}{E_{t-1}} \quad (1)$$

This measure puts capital inflows in relation to the risk taking capacity of the U.S. financial system (as captured by banking equity) and directly arises

in the context of the stylized model that we present in Section 4 to guide our bank-level analysis in Section 5. As our main measure of house prices at the state-level, we use the house price index compiled by Davis and Heathcote (2007) which is maintained at the Lincoln Institute of Land Policy. This data set has the major advantage that it corrects for differences in housing quality across time and states. It is therefore particularly suitable for interstate comparisons as we conduct them in this paper. Our main measure of state-level deregulation, denoted by $OPEN^s$, is the number of years that have passed until 1995 since state s first opened its banking market for out-of-state bank holding companies.³

Figure 1 documents the stylized fact that is the starting point for our analysis: the correlation between capital inflows and housing prices depends on a state's liberalization history. It plots $CAPFLOW_t$ and the four quarter ahead change of house prices for states that opened their banking markets early (by 1985) and those that liberalized late (after 1988) for the period from 1985 (the earliest date for which data on aggregate U.S. banking equity are available) to 2012.⁴

3. In the few cases where states maintained internal barriers to branching even after allowing in banks from outside the state, we use the minimum number of years that have passed until 1995 since either interstate banking or intra-state branching deregulation. This is because restrictions on intra-state branching would still effectively bar banks from outside the state to access all local markets, making it necessary to buy or open banks in all counties to be fully present in a state. Our results are not sensitive to this choice.

4. The two years 1985 and 1988 correspond to the lowest and highest quartile of deregulation dates respectively.

There is a striking correlation between the house price changes and capital inflows in early-deregulated states. This correlation is much weaker for the late deregulators. Interestingly, for the early deregulators house prices correlate with capital inflows already in the decade prior to 1995 whereas this correlation is virtually absent for the late deregulators. After 1995, the correlation is present for both groups of states, but the comovement between house prices and capital inflows remains much weaker for the late deregulators.

Our focus in this paper is on the period after 1997. This for at least three related reasons. First, during the 1980s and early 1990s, banking deregulation was ongoing in many states. Mian et al. (2017) show that the banking liberalization of the 1980s triggered a household credit boom that led to a boom and bust cycle in housing prices in states that had just opened while it did not affect those that were still relatively closed. This is exactly the pattern we see in Figure 1 for the pre-1997 period. By contrast, in the post-1997 period deregulation is completed and therefore pre-determined with respect to the surge in capital inflows from 1997 onward. Both groups of states now show a correlation between capital inflows and house prices, but the sensitivity of house prices among the late deregulators continues to be much lower. There seems to be a long shadow of deregulation.

Secondly, the emergence of large surpluses in China and other emerging economies from the mid-1980s onward suggests that at least a substantial part of U.S. capital inflows from the mid-1990s onwards can be explained by a shift in the global supply of capital. No such obvious exogenous shifter is apparent for the period before 1995, for which the results by Mian et al.

(2017) and others suggest that capital inflows into the U.S. rather were an endogenous reaction to domestic financial liberalization. Indeed, we show that the correlation between house prices and the current account vanishes during the pre-1997 period once we control for a range of domestic credit supply factors as well as for the state characteristics that Mian et al. (2017) have shown to be the vectors of the transmission of banking deregulation during the 1980s. By contrast, the correlation in the post-1997 period is robust to these controls.

Third, Value-at-Risk (VaR) models of risk management only started to be adopted widely by banks following the Basel committee's approval of the use of proprietary VaR measures for regulatory purposes in 1996 (Holton (2003)). Hence, the regulatory framework for banks led to the adoption of models that allowed banks to increase leverage directly as a function of their notional portfolio risk, including their geographical diversification. This mechanism is key for our empirical results.

While we do not wish to imply that the mechanism that we identify here could not have been present in the pre-1995 period, we argue that the post-1997 period provides a particularly suitable testing ground for it. Also, analyzing the extent to which capital inflows contributed to the rise of house prices prior to the biggest housing crisis in modern economic history is clearly of interest in its own right.

Table 1 presents summary statistics for our state-level data set while a detailed description of all data is available in Appendix A. For each variable and group of states (early or late deregulators) we compute the cross-state mean and median of the (within-state, time-series) mean as well as the average

of the state-level standard deviations. Note that early and late deregulators are surprisingly similar in terms of the average house price growth over the 1997-2012 period. However, the average volatility of early deregulators is much higher. This reflects the pattern familiar from Figure 1 that house prices in early deregulated states increased more during the boom period until around 2005 but also declined more strongly afterwards. The same pattern is also apparent once we turn to pre-sample trends (1985-1986) in the lower part of the table. Again, early and late deregulators had very similar house price and credit growth on average, but the average volatility for both variables was much higher among the early deregulators—consistent with the findings by Mian et al. (2017).

In Figure 2 we provide cross-sectional evidence on the long shadow that state-level liberalization in the 1980s cast on house price growth in the period of global imbalances. The upper panel of the figure plots average state-level housing prices in the pre-sample period (1985-1996) against $OPEN^s$. There is a mildly positive but insignificant relation. The second panel plots average house price growth for the period 1997-2005, when capital inflows surged, against $OPEN^s$. Now, the correlation is much stronger than in the pre-sample period, positive and significant. In the lower panel, we plot average house price growth over 2006-2012 against $OPEN^s$. During this period, when capital inflows declined sharply, there is a significantly negative correlation.

Figure 3 illustrates the link between capital inflows and house prices at the level of individual states for the period 1997-2012. For each state s , we run a

time-series regression of changes in house prices on capital inflows of the form

$$\text{HPGROWTH}_t^s = \alpha^s \times \text{CAPFLOW}_t + \beta^{s'} \times \text{CREDIT}_t + \delta_s + \epsilon_t^s, \quad (2)$$

where HPGROWTH_t^s is the four-quarter ahead growth rate in house prices in state s , CAPFLOW_t is our capital flow measure from (1) and CREDIT_t is a vector of aggregate credit supply and demand factors that we discuss in more detail below. Figure 3 plots the estimates of the coefficient α^s against our key measure of state-level openness—the numbers of years that elapsed between the date at which a state allowed the entry of out-of-state banks to its local banking market and 1995. The plot shows a clear positive relation: the correlation between house prices and capital inflows is higher in states that liberalized earlier. The cross-sectional regression of the estimated α^s on the number of years that have passed between liberalization until 1995 yields a positive and highly significant coefficient.⁵

3.2. Empirical framework and identification

Taking stock of the results in Figures 1, 2 and 3, our main state-level specification allows capital inflows into the U.S. load differently on different states as a function of a state's ex ante financial openness:

5. The picture would look much the same if we were to drop CREDIT_t from all regressions. While most of the estimated coefficients α^s are positive (consistent with our hypothesis), we note that individual values of α^s in Figure 3 may well be negative. The state-by-state time-series regressions (2) do not control for common time-variation across states nor do they account for the impact that capital inflows could have on house prices through a host of other state-level characteristics other than OPEN^s .

$$\text{HPGROWTH}_t^s = \alpha \times \text{OPEN}^s \times \text{CAPFLOW}_t + \text{CONTROLS}_t^s + \tau_t + \delta_s + \varepsilon_t^s \quad (3)$$

The dependent variable, HPGROWTH_t^s , is the four-quarter ahead growth in house prices in state s . On the right hand side of equation (3), OPEN^s is our measure of openness of the banking market in state s and CAPFLOW_t is again our measure of aggregate capital inflows. We also include time (τ_t) and state fixed effects (δ_s) along with a vector of controls, CONTROLS_t^s , that vary by time and location and that we discuss shortly.

Specification (3) interacts *aggregate* U.S. capital inflows as a common shifter of housing valuations with *ex ante* measures of state-level openness. From the outset, it therefore resolves many endogeneity issues that arise from the simultaneity of capital inflows and house prices at the local level. Any contemporaneous shock to the demand for and supply of housing (and credit) that is local-specific (and thus uncorrelated with country-wide capital inflows), will be absorbed by ε_t^s . Since $\text{cov}(\text{OPEN}^s \times \text{CAPFLOW}_t, \varepsilon_t^s \mid \text{CONTROLS}_t^s) = 0$, OLS estimates of regression (3) will correctly identify our parameter of interest, α .

Still, some challenges to identification remain. In particular, specification (3) does not allow us to infer whether capital flows are predominantly driven by global supply (a savings glut) or by an increased U.S.-wide demand for housing (and thus for credit). For example, if country-wide demand for mortgages loads more on open states—e.g. because integrated banks shift credit to states with high housing demand as recently shown by Loutskina

and Strahan (2015)—this could lead to higher house prices and higher inflows in open states without inflows being causal for house prices. Also, it could still be the case that some states liberalized earlier than others because of a latent demand for housing credit. For example, one may expect that political pressure to deregulate access to bank credit is higher in relatively densely populated states with many building restrictions. All of these examples would suggest that there is a higher demand for housing and mortgage credit in more open states, such that loan demand in state s is $L_t^{D,s} = \text{OPEN}^s \times L_t^D$ where L_t^D is some U.S.-wide credit demand factor. Since it is likely that $\text{cov}(\text{CAPFLOW}_t, L_t^D) > 0$, this would lead to an upward bias in our OLS estimates. In the same mold, other, domestic credit supply factors such as monetary policy could impact housing prices more in more open states, again leading us to overestimate the effect of capital inflows.⁶

It is well-documented in the literature that bank deregulation had considerable positive and persistent effects on income, output growth and many other state-level outcomes already in the short term (Kroszner and Strahan (2014)). Mian et al. (2017) show that deregulation led to a consumer credit boom during the 1980s and that it increased the importance of the non-tradable sector. If the impact of capital flows on house prices is heterogeneous

6. Favilukis et al. (2012) show that such domestic credit supply completely absorb the predictive power of capital inflows for house prices in aggregate time series regressions. Taylor (2007) and Borio and Zhu (2012) have emphasized the role of these factors in the context of the risk-taking channel and Jorda et al. (2015) provide long-run historical evidence for housing markets. Del Negro and Otrok (2007) document a limited role for monetary policy in driving the U.S. housing boom.

across states because of such differential pre-sample trends rather than because of differences in financial openness itself, we would also get biased estimates.

A third concern is that concurrent trends in financial liberalization in the years before the financial crisis affected house prices, rather than capital inflows themselves. As shown by Favara and Imbs (2015), interstate branching deregulation had a major impact on mortgage supply during our sample period, so it is clearly important to control for it, since it happened concurrently with the surge in capital inflows during the second half of the 1990s and the early 2000s.

To alleviate these concerns as best we can, we include three sets of controls in all our state-level regression specifications. First, we include a vector of state-time-varying variables consisting of income and population growth and the house price–income ratio. In particular, this vector also includes the time-varying interstate-branching deregulation indicator by Rice and Strahan (2010). We call these the baseline controls.

The second set of controls is the interaction of $OPEN^s$ with time-varying aggregate variables that affect the supply or demand of credit and that we collect in the vector $CREDIT_t$. This vector comprises of the domestic credit supply factors considered in Favilukis et al. (2012), i.e. the long-term (10 year) and short term (3 months T-Bill) real interest rate, the survey of attitudes of senior loan officers as well as the Baa-Aaa spread. We also construct a measure of deviations of the monetary policy rate from a Taylor rule. To gauge credit demand, we use data from the Michigan survey of Consumers used in Lambertini et al. (2013) about buying conditions for houses. This measure is

the difference between the fraction of surveyed households who think it is a good time to buy a home and the fraction that thinks it is not a good time. Taken together, these credit (demand and supply) factors should considerably mitigate the concern that our findings are due to omitted aggregate credit factors that could have impacted house prices more in financially more open states.

The third set of controls are interactions of capital flows with a host of pre-1997 state-level characteristics. While all our specifications of (3) control for state fixed effects, it may still be the case that different pre-1997 trends that are correlated with financial integration affect the post-1997 impact of capital inflows on house prices. We therefore include interactions of $CAPFLOW_t$ with 1985-1996 averages of all the state-level variables summarized in the lower panel of Table 1: house price growth, the growth rates of tradable and non-tradable GDP and wages, volatility of GDP growth, employment growth, the unemployment rate, real estate lending growth, exports relative to GDP (Michalski and Ors (2012) and a measure of import exposure (Autor et al. (2013)) aggregated to the state-level. We denote this vector of pre-1997 state-level characteristics by $PRE97$.

Finally, we conduct a host of robustness checks, including placebo assignments of regulation dates and running our baseline specification on the pre-1997 data set. In the second part of the paper, we then tackle any remaining identification issues by going granular, i.e by studying the impact of capital inflows directly at the bank level and then aggregating up banks' decisions to the county-level.

3.3. State-level results

Our analysis at the state level is based on data for the 47 contiguous U.S. states excluding Delaware (which is a financial center and therefore is special concerning banking regulation). A detailed description of the data is provided in Appendix A. Table 2 presents our baseline results for the period 1997–2012—the period of global imbalances—based on these quarterly state-level data. Consistent with our preliminary analysis in Figures 1 and 3, Table 2 shows that housing changes in more open states are significantly more exposed to aggregate capital inflows into the United States. This result also holds up once we add our credit variables interacted with openness, $OPEN^s \times CREDIT_t$, and the interaction of capital flows with pre-1997 characteristics, $CAPFLOW_t \times PRE97^s$.

In the last column, we also report results from a specification in which we interact our credit variables with state dummies and the pre-1997 characteristics with time effects. This lets $CREDIT_t$ load on house prices through arbitrary channels (instead of only) via $OPEN^s$ and allows for any time-varying factor to affect house prices via pre-1997 characteristics. Even now, our coefficient of interest remains stable and is still significant at the 15 percent level.

The coefficient α captures the annualized impact of capital inflows on house prices. In the specifications in Table 2, the estimate of α fluctuates around a value of 0.03. It is informative to discuss the implications of this estimate for the heterogeneity of house price developments across states as well for aggregate U.S. house prices. Measured relative to total banking equity,

capital inflows into the United States amounted to an average of 15 percent on a quarterly basis over the period 1997-2012 and of 17.5 percent over 1997-2005. According to Table 1, the median early deregulator opened its banking market 6 years earlier than the median late deregulator. This implies that, over the 8 years of the housing boom between 1997 and 2005, the median early deregulator would have had house prices that were $(1 + 0.175 * 0.03 * 6)^8 - 1 = 28\%$ percent higher than in the median state that deregulated late.

These results show that capital inflows had a very heterogeneous impact on house prices across states. But what do our results imply for aggregate U.S. house prices? Figure D.1 shows results for two scenarios: one in which the impact of capital inflows on the level of house prices is assumed to be permanent and one in which it is transitory, fading out over eight quarters.⁷ We can see from the figure that house prices for the average state virtually doubled between 1997 and 2007. Under the permanent-impact scenario, capital inflows would account for almost 60 percent of this run-up. Under the transitory-impact scenario, they still account for 40 percent. The range of these estimates is sensitive to the specific assumptions underlying each scenario. Their order of magnitude, however, suggests that capital inflows were an economically important factor behind the pre-crisis house price boom.

Robustness across sub-samples, placebo treatments. In Table 3 we document that our results are not driven by specific states. We first drop the so-called “sand states” (Arizona, California, Florida, Nevada) from the sample in which

7. Appendix D provides details on how we construct the aggregate house prices series for the two scenarios.

the housing booms before 2007 were particularly pronounced. We then drop some agricultural states from the sample that have few big cities—where house price increases were likely to be strongest—and that also were slow to deregulate (Montana, Kansas, North Dakota, Nebraska and Idaho). We also run a specification in which we drop South Dakota and Maine. Maine appears as an outlier in Figure 3 and as noted by Mian et al. (2017), South Dakota used the abolition of usury laws to attract credit card business. Finally, we exclude California and New York. Our identification assumption is that developments in individual states have no impact on aggregate U.S. capital inflows and this assumption could be violated for these two biggest and most populous states. Our results remain unchanged across all of these sub-samples.

We also examine the robustness of our results on the pre- (1997-2005) and post-crisis (2006-2012) subperiods separately. Splitting the sample in half while keeping all of our controls reduces the significance of the results to just below the 10 percent level, but our coefficient of interest remains stable across the two periods.

As a placebo test, we run our baseline regression on the pre-1997 sample. If our hypothesis of deregulation casting a long shadow is correct, then our coefficient of interest should be insignificant in the pre-1997 period. Indeed, our estimate of α collapses and becomes insignificant on the pre-sample period.

The differential impact of capital flows on house prices across states could be related to some unobserved trend that has been unfolding in a similar way to banking deregulation over time. To verify that this is not the case, we assign placebo measures of openness to each state by drawing from the empirical

cross-state distribution of $OPEN^s$. We then run our baseline regression on 10000 placebo sample draws. The results, in the last column of Table 3, show that the interaction between capital inflows and the placebo openness measure is zero on average. The standard deviation of the estimated coefficient across these 10000 draws is 0.013 and the individual coefficients are significant in less than 5 percent of the cases. Thus, the significance of our baseline regression on real-world data is clearly not spurious.

Alternative measures of capital inflows, net vs. gross flows. Appendix D provides evidence that our results are robust to alternative measures of capital inflows.

3.4. Are banks in early-liberalized states different?

Our state-level results raise the question how banks in early liberalized states differ from those in states that liberalized late. We focus on two dimensions of interstate heterogeneity of banks: their geographical diversification and the composition of their balance sheets, notably their equity ratios. These are the two dimensions of bank heterogeneity along which banks differ in their response to capital inflow shocks in our model in the next section.

The time passed since banking liberalization should be a very good proxy of the average geographical diversification of its banks: first, in states that have been open for longer, out-of-state banks had a longer time to establish themselves. Secondly, since interstate liberalization often took place on a mutual basis (i.e. banks were only allowed to enter if their home states allowed entry), banks headquartered in early-liberalizing states had more

opportunities to diversify to other states. Both effects should lead to higher local market shares of banks with a high level of diversification.

We also expect geographically more diversified banks to have a higher risk taking capacity, which should be reflected in lower equity ratios (higher leverage). One may also conjecture that geographically more diversified banks have easier access to wholesale funding from nationwide or global capital markets.

The evidence presented in Table 4 and Figures D.2 and D.3 supports these hypotheses. We measure the degree of geographical diversification of banks in a state in four different ways: i) the interstate asset ratio proposed by Morgan et al. (2004), ii) the interstate mortgage ratio, iii) the average number of states in which a state's banks are active, and iv) the average (bank-level) geographical concentration of bank assets in a state.⁸ For all four indicators we find a significantly positive relationship with the time elapsed since banking deregulation. We also find that the market-share-weighted average equity ratio of a state's banking system is significantly lower for early liberalizers. Finally, the data also support the view that banks in early-liberalized states finance a higher share of their balance sheet with wholesale funding and a lower share with retail deposits.

As a further robustness check, in Table D.1, we perform versions of our baseline regressions in which we interpret these measures of geographical diversification as de facto measures of a state's degree of banking integration

8. We provide details of the construction of these and the other indicators of banks' characteristics used in this subsection in Appendix D.

and use them instead of our *de iure* measure $OPEN^S$. Our earlier results remain: capital inflows load more strongly on house prices in states with more integrated banking sectors (geographically more diversified banks).

In Appendix B we show formally that the state-level results in this section can be interpreted as the result of the aggregation of banks' heterogeneous credit-supply responses to a positive aggregate refinancing shock—the savings glut. Appendix B extends the reduced-form model by Landier et al. (2017) by allowing banks to differ in their exposure to aggregate capital inflows. If more exposed banks extend credit more and have a higher market share in states that deregulated early, the impact of capital inflows on credit supply in early-deregulated states will be stronger and house prices will increase more. Sections 5.1 and 5.4 below will provide bank-level evidence that supports the prediction of our theory that geographically diversified banks expanded credit more and Section 5.3 shows that this did indeed drive up house prices.

4. A theoretical model of the transmission mechanism

We have established that global imbalances have a stronger bearing on house prices in states with more integrated banking markets. As we have seen in the previous sub-section, states that liberalized earlier have a higher presence of geographically diversified banks and these banks have higher leverage. In this section, we examine the transmission mechanism between capital flows and house prices in more detail. We propose a simple model in which we interpret the savings glut as a positive refinancing shock that, *a priori*,

affects all banks equally. However, due to their lower notional balance sheet risk, geographically diversified (“integrated”) banks are better placed to take advantage of this refinancing shock to expand their lending. This can explain why house prices increased more in states that liberalized early and therefore have a strong presence of geographically diversified banks.

The model builds on Shin (2012) and assumes that banks are risk neutral but face a value-at-risk (VaR) constraint that is imposed by the regulator. The value at risk constraint stipulates that the bank maintain a constant probability of default. We present details of the setup of the model and the derivation of the main equations in Appendix C. A key implication of the model is that – as in Shin (2012) – the VaR constraint leads the bank to implement a constant leverage given the risk and expected excess return on its loan portfolio. Specifically, the bank’s leverage ratio can be written as

$$\text{LVG}_t \equiv \frac{A_t}{E_t} = \frac{(1 + r_t^d)}{\Phi \sigma_n - E(r_{t+1}^l - r_t^d)} \quad (4)$$

where A_t is the value of the banks’ total assets (that we assume is composed only of mortgage loans, L_t) and E_t its equity. The banks’ lending activities generate a return r_{t+1}^l while the bank pays its lenders a deposit rate r_t^d . The risk of the banks’ portfolio is measured by the standard deviation of the return, r_{t+1}^l which we denote by σ_n to indicate that it will decline with the number of locations (e.g. states) n in which a bank is doing business so that $\sigma_{n_1} < \sigma_{n_2}$ for $n_1 > n_2$. The parameter Φ measures the bank’s (constant) distance to default expressed here in multiples of the standard deviation of its loan portfolio.

Now consider how leverage is affected by a banks' degree of regional diversification. To this end, we compare an integrated bank I to a local bank L . Bank I is geographically diversified and operates in $n > 1$ states with low portfolio risk $\sigma^I = \sigma_n$. The local bank L operates only in $n = 1$ states and thus has a high portfolio risk $\sigma^L = \sigma_1$, so that $\sigma^I < \sigma^L$. It is clear from equation (4) that the more diversified bank should have higher leverage *ceteris paribus*. It is also easy to see that leverage should increase for both banks as r^d drops. Importantly, however, the leverage of the I -bank is more sensitive to a drop in the refinancing rate r^d than leverage of the L -bank, $\partial \text{LVG}^I / \partial r^d > \partial \text{LVG}^L / \partial r^d$. Keeping equity constant, this implies that a drop in the refinancing rate r^d will lead to larger increase in lending for the integrated bank than for the local bank.

We follow Caballero and Krishnamurthy (2009) and model the savings glut as an exogenous increase in the supply of funds deposited into the U.S. banking system. These deposits are safe from the point of view of the international investor. Keeping the degree of regional banking integration fixed—an assumption that we will relax below—equilibrium in the market for deposits requires

$$\frac{\Delta D_{t+1}^{US}}{E_t^I + E_t^L} = \left(\frac{\partial \text{LVG}^I}{\partial r^d} \lambda_t + \frac{\partial \text{LVG}^L}{\partial r^d} (1 - \lambda_t) \right) \Delta r_{t+1}^d \quad (5)$$

where ΔD_{t+1}^{US} denotes the change in funds deposited in the U.S. banking system, E^I and E^L denote the equity of integrated and local banks respectively, and $\lambda_t = E_t^I / (E_t^I + E_t^L)$ is the share of total U.S. banking equity that is invested in integrated banks. We can think of $\Delta D_{t+1}^{US} / (E_t^I + E_t^L)$ as the model

counterpart of our capital inflow measures and λ_t as an indicator of regional banking integration. The model then implies that the global demand for U.S. safe assets (i.e. deposits into the banking system) leads to a decline in the refinancing rate for U.S. banks. This decline just relaxes banks' VaR constraints sufficiently to allow them to absorb the additional foreign demand for U.S. safe assets. Since integrated banks' leverage is more sensitive to the drop in refinancing rates (i.e. $\partial \text{LVG}^I / \partial r^d > \partial \text{LVG}^L / \partial r^d$) they end up absorbing a larger share of these capital inflows, thus increasing lending more than local banks. By implication we expect to see that, for a given drop in the refinancing rate, aggregate lending increases more strongly in states where integrated banks have a higher market share to start with.⁹

Gross versus net flows. So far, we have modeled the shock to foreign demand for U.S. bank deposits—the 'savings glut'—as an exogenous increase in gross capital inflows, $\Delta D_{t+1}/E_t$. However, we are interested in how these capital inflows affect *domestic* lending for a *domestic* end use (i.e. residential housing). If U.S. banks absorb a part of the savings glut to increase foreign lending or if foreign banks take advantage of cheap dollar funding conditions to increase lending in their home countries, this will dampen the required adjustment in domestic lending. Once we introduce a foreign bank in our

9. We assume that adjustments in leverage take place exclusively via adjustments in debt, i.e. equity is not actively managed by the bank. Adrian and Shin (2010b) show that this is a very good empirical characterization of the actual behavior of U.S. commercial banks. Under this assumption, variation in leverage directly translates into variation in loan supply, since banks' lending supply L_t is given by $L_t = A_t = \text{LVG}_t \times E_t$. Shocks to equity then only act as exogenous shifters of the loan supply function.

model above, the equilibrium condition in the U.S. deposit market becomes

$$\frac{\Delta D_{t+1}^{US}}{E_t^I + E_t^L + E_t^F} = \underbrace{\left(\frac{\partial \text{LVG}^I}{\partial r^d} \lambda_t + \frac{\partial \text{LVG}^L}{\partial r^d} (1 - \lambda_t) \right) \frac{E_t^I + E_t^L}{E_t} \Delta r_{t+1}^d}_{\text{increase in domestic lending}} + \underbrace{\frac{E_t^F}{E_t} \frac{\partial \text{LVG}^I}{\partial r^d} \Delta r_{t+1}^d}_{\text{capital outflows}}$$

where $E_t = E_t^I + E_t^L + E_t^F$ is again total U.S. banking equity. The first term on the right hand side now reflects the increase in domestic lending while the second gives capital outflows (i.e. the increase in foreign lending). Given total banking equity E_t , the presence of foreign banks (or of U.S. banks committing a part of their equity to foreign operations) dampens the upward pressure on domestic lending. Again using our assumption that changes in leverage only come from changes in lending, we see that *net* capital inflows relative to bank equity become the relevant statistics for the increase in domestic lending:

$$\text{CAPFLOW}_{t+1} := \frac{\Delta D_{t+1}^{US} - \Delta L_{t+1}^F}{E_t} = \left(\frac{\partial \text{LVG}^I}{\partial r^d} \lambda_t + \frac{\partial \text{LVG}^L}{\partial r^d} (1 - \lambda_t) \right) \frac{E_t^I + E_t^L}{E_t} \Delta r_{t+1}^d \quad (6)$$

The left hand side of (6) is exactly the measure we use as our baseline. However, we also re-run most specifications in the paper with gross inflows. The findings are very similar, but generally a little less strong. This is consistent with our theory: if the propensity to “recycle” inflows into foreign lending (given by E^F/E) is approximately constant, then gross inflows should give us broadly the same picture. However, net inflows should identify the impact on domestic credit growth more sharply if E^F/E moves over time.

Loan supply and demand shocks. Figure D.4 illustrates the impact of liquidity supply and demand shocks on refinancing and lending rates. A savings glut shock will drive down banks’ refinancing and lending rates. while a loan

demand shock should drive them higher. In our bank-level analysis below we document that geographically diversified banks lowered their interest rates in response to capital inflows, consistent with the supply-shock scenario.

The banking glut. Appendix D also provides a discussion of how our framework and empirical analysis relate to the increase in international gross banking positions in the run-up to the financial crisis—the banking glut documented by Shin (2012).

5. Evidence from bank-county-level data

We now turn to examining the empirical predictions of our model. In so doing, we face two empirical challenges. First, we need data that allow us to identify banks' extent of geographical diversification. Second, we need to isolate the effect of banks' credit supply from local demand factors in such data.

We address the first of these challenges by turning to the Home Mortgage Disclosure Act data base (HMDA). The HMDA collects all mortgage applications by private households in the United States handled by financial institutions that exceed a low, annually adjusted threshold in mortgage assets. It contains detailed information about the geographical location of the property, whether the application was successful (i.e. the mortgage loan was granted) and the loan size. It also provides information about the financial institution that processed the loan, such as its regulatory ID number, the institution type (commercial bank, credit or savings union, mortgage company), total assets as well as the institution's home state. We aggregate the HMDA data to obtain the annual growth rates of mortgage lending by

institution and county for the period 1997-2005. We focus on around 1800 counties in the United States that are part of a core based (metropolitan or micropolitan) statistical area (CBSA).

We then consider the following decomposition of the growth rate of mortgage lending by bank b in county c :

$$\frac{L_t^{b,c} - L_{t-1}^{b,c}}{L_{t-1}^{b,c}} = \beta_t^b + \gamma_t^c + \nu_t^{b,c} \quad (7)$$

where β_t^b is a bank-time-specific effect and γ_t^c is a county-time effect and $\nu_t^{b,c}$ a residual that we assume is i.i.d across banks and counties as well as over time. Since β_t^b varies only by bank, it is free of location-specific demand factors and we can think of the bank-time effect as isolating the supply component of mortgage lending of bank b that is common across all counties. In the same spirit, since γ_t^c varies only by county, we can think of it as reflecting the local-specific demand component that all banks face in a given county c .

To obtain economically meaningful bank-specific supply factors, we propose to adapt the method proposed by Amiti and Weinstein (2018) for matched bank-firm data to our setting here, where counties that borrow from different banks take the place of firms. Two features of the method proposed by Amiti and Weinstein (2018) make it particularly suitable for our purpose.

First, the method yields estimates of β_t^b and γ_t^c that are consistent with aggregation. It therefore explicitly takes account of the granular structure of

the banking sector which is characterized by a few relatively large banks dominating many local markets.¹⁰

To see the gist of the methodology, assume there are B banks and C counties and let $\omega_{t-1}^c = [\omega_{t-1}^{1,c}, \dots, \omega_{t-1}^{b,c}, \dots, \omega_{t-1}^{B,c}]'$ the vector of market shares of all banks in county c at time $t-1$ and $\phi_{t-1}^b = [\phi_{t-1}^{b,1}, \dots, \phi_{t-1}^{b,c}, \dots, \phi_{t-1}^{b,C}]'$ the vector of shares of each county in bank b 's country-wide mortgage portfolio. Let us further collect the bank-time and county-time effects into the vectors β_t and γ_t respectively and let Λ_t be the $B \times C$ matrix of bank-county-level growth rates with $\Lambda_t^{b:}$ its b -th row and $\Lambda_t^{:c}$ its c -th column. Then, the growth rate of bank b 's mortgage lending can be written as

$$\frac{L_t^b - L_{t-1}^b}{L_{t-1}^b} = \Lambda_t^{b:} \phi_{t-1}^b = \beta_t^b + \phi_{t-1}^{b'} \gamma_t \quad (8)$$

and the growth rate of lending for each county c is

$$\frac{L_t^c - L_{t-1}^c}{L_{t-1}^c} = \omega_{t-1}^{c'} \Lambda_t^{:c} = \gamma_t^c + \omega_{t-1}^{c'} \beta_t \quad (9)$$

where we have made use of the fact that the elements of ϕ_{t-1}^b and ω_{t-1}^c sum to one and where the i.i.d.-ness of $\nu_t^{b,c}$ ensures that $\sum_{b=1}^B \omega^{b,c} \nu_t^{b,c}$ and $\sum_{c=1}^C \phi^{b,c} \nu_t^{b,c}$ go to zero for sufficiently large B and C . Amiti and Weinstein show that the two sets of moment conditions (8) and (9) uniquely identify (up to normalization) the vectors β_t and γ_t . Hence, there is only one solution to

10. Landier et al. (2017) explore the role of banking sector granularity for house price synchronization across U.S. states. Bremus et al. (2018) show that shocks to big banks have a disproportionate effect on macroeconomic outcomes at the international level.

the decomposition (7) that satisfies the condition that the (market or portfolio share) weighted sums of bank and county effects sum to bank- and county-wide aggregates respectively.

The second advantage of the method by Amiti and Weinstein is that it allows to deal with the creation of new lending relationships between banks and counties.¹¹ As we will discuss in more detail below, our theory implies that a positive funding shock for banks will increase their incentive for geographical diversification. Furthermore, Favara and Imbs (2015) have shown that the geographical expansion of banks after the wave of inter-state branching deregulation during the second half of the 1990s and the early 2000s had a major impact on banks' credit supply. Hence, our theory and prior empirical research by others both suggest that the extensive margin of lending is important in our sample. These considerations make the approach by Amiti and Weinstein (2018) our method of choice.

5.1. *Bank-level results*

Solving the moment conditions (8) and (9) gives us estimates of the institution-level supply shock β_t^b for all mortgage lenders in the HMDA and loan demand

11. Amiti and Weinstein show that the weighted-least squares estimate of the decomposition (7) with weights $L_{t-1}^{b,c}$ satisfies the moment conditions (8) and (9) if and only if there is no new lending along the extensive margin, i.e. if banks only lend to counties to which they already had a lending relationship. In the general case when there is new lending along the extensive margin, the moment conditions have to be solved explicitly for β_t and γ_t , which requires the solution of a $(B + C - 2)$ -dimensional system of linear equations for each period t . This is the approach we follow in our analysis here.

shocks γ_t^c for all counties part of a CBSA. For our subsequent analysis it is, however, informative to distinguish between two types of mortgage lenders in the HMDA: insured depository institutions (banks) and non-depository institutions (mortgage companies). Demyanyk and Loutskina (2016), discuss that mortgage companies are exempt from the specific regulation that is meant to curb the risk-taking behavior of insured depository institutions. For this reason and because mortgage companies primarily originate loans to sell them for securitization, the model of banks' risk taking behavior in the previous section should not apply to them. Mortgage companies therefore provide a natural placebo sample that allows us to check that our mechanism is really specific to banks.

Bearing this in mind, we now turn to testing the implications of our theory. It is straightforward to extend the model in the previous section to many banks. The change in aggregate leverage can then be written as

$$\Delta LVG_t = \left[\sum_{b=1}^B \frac{\partial LVG_{t-1}^b}{\partial r} \lambda_{t-1}^b \right] \Delta r_t^d$$

Adrian and Shin (2010a) show that, prior to the 2008 financial crisis, banks predominantly used the asset side of their balance sheet (i.e. in our setting lending) to target their leverage, treating equity in a passive way. This implies that changes in leverage are (mainly) driven by changes in lending. With equity kept constant, the leverage response of bank b in the expression above can be written as

$$\frac{\partial LVG_{t-1}^b}{\partial r} \lambda_{t-1}^b = \frac{\partial L_{t-1}^b / \partial r}{L_{t-1}^b} \times \frac{L_{t-1}^b}{L_{t-1}} \times LVG_t$$

Using the expression for the bank's leverage in (4), we then obtain the contribution of bank b to aggregate credit growth:

$$\frac{\partial L_{t-1}^b / \partial r}{L_{t-1}^b} \times \frac{L_{t-1}^b}{L_{t-1}} \Delta r^d = \underbrace{\left[\frac{1}{\Phi \sigma_{t-1}^b - E(r^l - r^d)} - \frac{1}{1 + r^d} \right]}_{b's \text{ marginal risk taking capacity (RTC}_{t-1}^b)} \times \frac{L_{t-1}^b}{L_{t-1}} \left(-\Delta r_t^d \right) \quad (10)$$

We interpret the term in rectangular brackets as the marginal risk-taking capacity (RTC_{t-1}^b) of the bank in response to a decline in the bank's refinancing rate, Δr^d . Note that cross-bank variation in RTC depends only on two bank-specific variables: the bank's portfolio risk, σ^b , and its share in aggregate lending, L^b/L . The marginal risk taking capacity increases as the bank's notional portfolio risk σ^b decreases and it is proportional to the bank's share in aggregate lending.

We now correlate our bank-level shocks β_t^b with $\text{RTC}_{t-1}^b \times (-\Delta r)$. The interest rate decline in our model is caused by the capital inflow. Hence, we can write $-\Delta r = \alpha \text{CAPFLOW}$ for $\alpha > 0$ and we therefore expect a positive coefficient α in a regression of the form

$$\beta_t^b = \alpha \times \text{RTC}_{t-1}^b \times \text{CAPFLOW}_t + \text{CONTROLS} \quad (11)$$

where the vector of controls contains bank-specific variables such as total assets or the bank's aggregate asset growth rate as well as interactions of RTC with other economy-wide credit supply factors, as we have included them in our state-level regressions before.

To compute RTC, we measure σ^b as the product of the Herfindahl index of the geographical concentration of the bank's mortgage portfolio with the

average annual standard deviation of house price growth

$$\sigma_{t-1}^b = \overline{\text{SD}(\Delta\text{HPGROWTH}_l)} \times \sqrt{\sum_{l \in \mathcal{L}(b)} (\phi_{t-1}^{b,l})^2} \quad (12)$$

where where $\mathcal{L}(b)$ denotes the set of locations (counties, MSAs or states) in which bank b is active and $\phi_{t-1}^{b,l}$ again denotes the share of the bank's mortgage portfolio held in location l . The term $\overline{\text{SD}(\Delta\text{HPGROWTH}_l)}$ denotes the the cross-sectional mean of the time-series standard deviations of house price growth at the location-level.

One choice we face here is at which level of aggregation to compute this Herfindahl index to obtain a suitable measure of geographical diversification. Our data are at the bank-county level, so computing σ_{t-1}^b directly at the county-level appears as a natural starting point. But this could be misleading because the shocks affecting individual counties are likely to be correlated within an MSA or even within the state. We choose a pragmatic approach and provide our results for measures σ_{t-1}^b computed at the county-, MSA-, and state-levels. Finally, we assume that $E(r^l - r^d) = r^d = 0.01$ and set the distance to default $\Phi = 2.32$. Assuming that banks' returns are normally distributed, in our annual data set this amounts to an annual probability of default of 1 percent.¹²

12. Our choice of parameters is disciplined by the constraint that the notional bank leverage in equation (4) is positive and bigger than unity for any given empirical values of σ_{t-1}^b . Provided that this (economically sensible) constraint is fulfilled, our results are not sensitive to changes in the parameters.

To account for the granular structure of the banking market and to remain consistent with the adding up-constraint discussed in the context of the Amiti-Weinstein procedure above we estimate regression (11) by weighted least-squares (WLS) with the weights being given by a bank’s share in aggregate (U.S.-wide) lending in the previous period. All regressions include the credit-supply factors by Favilukis et al. (2012) also used in our state-level regressions (interacted with the respective RTC measure) and the level and growth rate of the banks’ total assets. We also include bank- and time-fixed effects in all regressions which allows us to capture any banking-sector specific shocks that are common to all banks as well as unobserved heterogeneity across banks, e.g. in terms of banks’ different business models.

Table 5 presents the results for banks (i.e. for depository institutions). Except for the case in which σ^b is computed at the county-level—likely for the reasons discussed above—there is a significantly positive relation between $\text{RTC}_{t-1}^b \times \text{CAPFLOW}_t$ and the bank-level shocks β_t^b .

These regressions provide strong evidence in favor of our conjectured mechanism. But they do not provide an estimate of the marginal contribution of capital inflows to bank lending. Ultimately, we want to isolate the component of bank-level shocks that we can ascribe to capital inflows alone—i.e. after controlling for all other credit supply factors and bank-specific controls. To construct such a bank-specific capital inflow shock, we re-run regression (11), but only with the controls, i.e. dropping the term $\text{RTC}_{t-1} \times \text{CAPFLOW}$. By construction, the residual ∇_t^b of this regression is orthogonal to the aggregate credit-supply factors (interacted with RTC) as well as to the purely bank-specific controls. In a second stage, we then run the

regression

$$\nabla_t^b = \alpha_1 \text{RTC}_{t-1}^b \times \text{CAPFLOW}_t + \alpha_2 \text{RTC}_{t-1}^b + \text{bank and time effects} \quad (13)$$

The fitted value $\hat{\nabla}_t^b = \alpha_1 \times \text{RTC}_{t-1}^b [\text{CAPFLOW}_t + \alpha_2 / \alpha_1]$ can then be interpreted as that part of the bank-level shock β_t^b that is due only to capital inflows.

The last three columns of Table 5 provide the estimates of regression (13), again for the three bank portfolio risk measures computed at the county, MSA and state levels respectively. We have a strongly significant relation between capital inflows and the part of the bank-level shock that remains unexplained by other economy-wide and bank-specific factors. This result holds irrespective of the level of geographical granularity at which we compute the degree of portfolio diversification of the bank.

Results on non-bank placebo sample. In Table D.4, we repeat the bank level analysis of Table 5 on our placebo sample of mortgage companies. As discussed above, our model should not apply to mortgage companies. This is what we observe in Table D.4: the coefficients on the interaction of capital inflows and RTC is small and insignificant, irrespective of the level of granularity at which we measure geographical diversification. Hence, geographical diversification has no bearing on how the lending supply of mortgage banks reacts to capital inflows.

Gross vs. net inflows. One prediction of our model is that—while gross inflows matter for how much total lending by U.S.-based banks should increase—net inflows are more informative when our interest is in the impact of capital flows on domestic lending for domestic end uses (as is clearly the

case for mortgage lending). Table D.5 illustrates this point. When we estimate the regression of bank level effects on gross inflows and our other credit factors, gross inflows are generally found to be significant, as was the case for net inflows. However, once we re-estimate this regression with inflows dropped and then regress the residual on the interaction of gross inflows with RTC as in (13), the coefficient on gross inflows alone remains positive but is now imprecisely estimated. When both gross inflows and gross outflows are used independently in the same regression in the residual-regression, we find that the coefficients are both of similar magnitude, insignificant and of opposite sign. Indeed, an F-test cannot reject that the two coefficients sum to zero. Hence, using gross inflows generally gives the same picture as net inflows, but the effects are less precisely estimated—consistent with our theoretical model in the previous section.

5.2. *Constructing county-level capital inflow shocks*

We have established that a significant component of bank's lending supply can be explained by capital inflows and that this component increases in banks geographical diversification. In a next step, we construct county-level exposures to bank-level capital inflow shocks as the weighted average of $\hat{\nabla}_t^b$ across all banks in a county:

$$\text{CAPFLOW}_t^c = \sum_{b \in B_{t-1}} \omega_{t-1}^{b,c} \times \hat{\nabla}_t^b \quad (14)$$

where B_{t-1} is the set of banks active (in any county) at time $t - 1$. A key decision we face here is how to choose the weights $\omega_{t-1}^{b,c}$. One apparently

obvious choice would be banks' market shares. But cross-county variation in bank's de facto market shares could itself be the outcome of differences in local credit demand.

To deal with this challenge, we build on our analysis in the first part of the paper. Specifically, we suggest a de iure measure of banks' market shares that is based on the regulatory history of banks' home state and of the states in which they are active. This measure makes use of the fact that the liberalization of a state's banking market in most cases amounted to a liberalization on a reciprocal basis.¹³ Our openness indicator OPEN^s used in our state-level analysis therefore generally also defines the year from which it became possible for banks headquartered in state s to access markets in other states – provided these states had already opened their markets themselves. We therefore define the de iure market share of bank b in county c as follows

$$\omega_{t-1}^{b,c} = \frac{\min \left\{ \text{OPEN}^{s(c)}, \text{OPEN}^{s(b)} \right\} \times \mathbf{1}_{\{s(c) \neq s(b)\}} + \mathbf{1}_{\{s(c) = s(b)\}} \times d(c)}{\sum_{b \in \mathcal{B}_{t-1}(c)} \left\{ \min \left\{ \text{OPEN}^{s(c)}, \text{OPEN}^{s(b)} \right\} \times \mathbf{1}_{\{s(c) \neq s(b)\}} + \mathbf{1}_{\{s(c) = s(b)\}} \times d(c) \right\}} \quad (15)$$

where $s(c)$ denotes the state to which county c belongs and $s(b)$ denotes the home state of bank b and $\mathcal{B}_{t-1}(c)$ is the set of banks active in location c in period $t - 1$. The numerator of this expression is the number of years (until 1995) that the state $s(c)$ in which bank b is active has been open for banks from bank b 's home state. For banks from outside the state, i.e. for which the

13. These mutual reciprocal liberalizations were initially often limited to states in the region and later expanded to the national level. We do not attempt to account for the complexity of these regional agreements here.

indicator functions are $\mathbf{1}_{\{s(l) \neq s(b)\}} = 1$ and $\mathbf{1}_{\{s(l) = s(b)\}} = 0$, this number takes the value $\text{OPEN}^{s(l)}$ if the home state of the bank liberalized earlier than state $s(c)$. For out-of-state banks from states that liberalized later, it takes the value $\text{OPEN}^{s(b)}$. For local banks for which state $s(c)$ is the home state ($\mathbf{1}_{\{s(c) \neq s(b)\}} = 0$, $\mathbf{1}_{\{s(c) = s(b)\}} = 1$) it takes on the value $d(c)$.¹⁴

The numerator therefore varies for each bank in state $s(c)$ as a function of the bank's home state. The denominator just sums this bank-specific number across all banks active in county c which ensures that $\sum_{b \in B_{t-1}(l)} \omega_{t-1}^{b,c} = 1$.

We can think of $\omega_{t-1}^{b,c}$ as the hypothetical market share that bank b should have in county c if banks' started to enter the target market c at the earliest possible date and then grow their lending at a rate that is equal across all banks. For each bank, this 'as-if' share only depends on its own home state, the regulation of the state to which the target market belongs as well as on the regulation of the home states of the other banks that are active in c . Our identifying assumption is that $\omega_{t-1}^{b,c}$ is uncorrelated with local demand conditions. We therefore use the weights from (15) in (14) to construct CAPFLOW_t^c .

14. To code this value, we assume that local banks could enter other local markets from 1980 inwards. Thus, we assume that banks in state s could generally access other local markets in state s before out-of-state banks because most state-liberalizations took place after 1980. As a robustness check, we also coded $d(c)$ as the years until 1995 since the dismantling of *intra-state* branching restrictions. In particular in states with strict unit banking, it could be argued that intra-state branching restrictions effectively barred entry of banks into particular counties even if these banks resided in the same state. This alternative coding did not materially affect our results, though.

5.3. *County-level results*

State-level average prices could mask heterogeneity within states. If housing supply elasticities differ across locations (e.g. due to different building and zoning laws or differences in physical terrain that affect the ease with which land can be developed), the savings glut could have differential effects on housing prices for reasons that are unrelated to a state's financial openness. In addition, even though we have included a battery of controls in our state-level regressions, there could still be unobserved state-level variables that co-vary with a state's liberalization history that could confound the causal interpretation of our state-level results. Having computed county-level capital inflow shocks that are constructed as aggregates of shocks at the disaggregate (bank) level, we can now address these remaining identification issues. Specifically, since $CAPFLOW_t^c$ is a complex function not only of the liberalization history of an individual target state but also of the liberalization histories of the home states of the banks operating in a given county, it is very unlikely to be correlated with any unobserved state-or county-level variables.

For our analysis in this subsection, we merge our county-level shocks with the county-level data set provided by Favara and Imbs (2015). This data set contains house prices and a rich set of controls for urban counties in the U.S. for the period 1995-2005. Specifically, the data set also includes a measure of housing supply elasticities at the county-level based on Saiz (2010). Table 6 shows county-level regressions of house prices growth on our county-level

capital inflow shocks of the general form

$$\text{HPGROWTH}_t^c = \alpha_0 \times \text{CAPFLOW}_t^c + \alpha_1 \times \text{ELASTICITY}^c \times \text{CAPFLOW}_t^c + \text{CONTROLS}$$

where ELASTICITY^c is the elasticity of housing supply in county c .

The first regression has no controls except county and time fixed effects. In the second column we add the set of controls considered by Favara and Imbs (2015): current and lagged income and population growth and a measure of the concentration of the local banking market (the Herfindahl index of the market shares of the local banks). We also add the interstate branching deregulation indicator compiled by Rice and Strahan (2010). In addition, we include the estimate of the county-level mortgage demand shock γ_t^c from (7). In column 3 we interact CAPFLOW_t^c as well as the interstate branching indicator with the housing supply elasticity to allow for the possibility that house prices might react less to a given credit supply shock if supply is very elastic. We also add lagged house price growth. This somewhat reduces the magnitude of our coefficient of interest but it remains significant.

In column 4, we add a full set of state-year dummies. Under this very demanding specification, the coefficient on the stand-alone term CAPFLOW_t^c is no longer significant, even though the estimate remains stable. It is now the interaction term $\text{ELASTICITY}^c \times \text{CAPFLOW}_t^c$ that is negative and significant. Hence, once all state-time varying factors are accounted for, the average effect of CAPFLOW_t^c on house prices becomes hard to identify. CAPFLOW_t^c still

matters—but its impact on house prices depends on local supply conditions and is stronger in inelastic markets.¹⁵

Our results also remain intact on a subset of counties that straddle state borders (column 5). We emphasize this point because we would expect that borrowers from other states within the same MSA could in principle take advantage of the capital inflow shock in a county in the neighboring state. This would tend to break the link between county-level capital inflows and house prices. While the estimated coefficient is indeed lower, it remains significant—even in the presence of the interstate-branching deregulation indicator by Rice and Strahan (2010).

Including a lagged endogenous variable in panel regressions is known to create a bias on the coefficient of the lagged endogenous variable that can also spill over into other coefficients of interest (Nickell bias). We therefore also estimate a specification of our regression using the GMM method by Blundell and Bond (1998). Again the results are robust.¹⁶

Table 7 examines to what extent our capital inflow shocks drive local mortgage lending. The first column considers total mortgage growth by all institutions. In the second column, we use the growth rate of mortgage lending by banks while in the third we focus on mortgage growth by mortgage

15. In particular, state-year effects absorb the interstate branching deregulation indicator as well as our state-level measure $OPEN^s \times CAPFLOW_t$.

16. As discussed in the online appendix of Favara and Imbs (2015), GMM could be very misleading if there is cross-sectional correlation in the errors. Therefore, we stick to our least-squares estimation as the baseline.

companies. The capital inflow shock is significant for county-level mortgage growth and, in particular, for the growth of mortgage lending by banks. It is not significant on the placebo sample, i.e. for the growth of non-bank mortgage lending.

Impact on lending growth and IV results. All of these results suggest that capital inflows had a causal effect on bank lending. This effect was stronger in local banking markets in which geographically diversified banks had more time to establish a strong foothold after the liberalization of state banking markets. We further strengthen this causal interpretation in columns 4 and 5 of Table 7. In column 4, we report a regression of housing growth on lending growth, using $CAPFLOW_t^c$ as instrument. In column 5, we again acknowledge that the ultimate impact of lending supply on house prices also depends on the housing supply elasticity. We therefore instrument lending growth and its interaction with $ELASTICITY^c$ using $CAPFLOW_t^c$ and $CAPFLOW_t^c \times ELASTICITY^c$.

In both specifications, the second-stage coefficients on lending growth are strongly significant. Turning to the first stages, in the specification with only one endogenous variable in column 4, the first stage F-statistics is well above the critical threshold of 10 (Stock et al. (2002)). In the specification in column 5, we have two endogenous variables and the simple first-stage F-statistics is not applicable. Instead, we report the Sanderson and Windmeijer (2016) conditional F-test of weak identification for each of the two variables. The test clearly rejects the null in both cases.

Endogenous geographical diversification: interstate branching and securitization. We emphasize that the wave of interstate liberalization during the 1980s that we have exploited in our empirical analysis so far allowed bank holding

companies from other states to acquire local banks. But acquired banks had to remain separate legal entities. Only later, and concurrently with the capital inflows hitting the U.S. from the second half of the 1990s onwards, the U.S. banking sector saw a second major wave of state-level financial liberalization: the gradual dismantling of remaining interstate bank-branching restrictions.¹⁷ Since this wave of liberalizations could be an important confounding factor, we have included the indicator of interstate-branching deregulation by Rice and Strahan (2010) and Favara and Imbs (2015) as a control variable in all our county- and state-level regressions in the paper. Our results show that capital inflows and a state's deregulation history have a strong impact on housing prices independently of the relaxation of inter-state branching restrictions.

In this section, we now go one step further. We ask to what extent both our county-level inflows shocks, as well as inter-state branching deregulation are driven by the interaction of capital inflows and a state's liberalization history. In our analysis, we directly build on Rice and Strahan (2010) who have argued that states with a strong presence of big, nationwide banks also saw the most forceful political lobbying for liberalization and eventually an earlier and more complete relaxation of restrictions. Branching liberalization in the 1990s could therefore have been foreshadowed by interstate banking deregulation during the 1980s in the sense that integrated banks could have lobbied for branching deregulation more successfully in states where they had

17. While the Riegle-Neal Act already stipulated that interstate branching restrictions had to be dismantled by 1995, states could opt out from this legislation and most did so, thus maintaining barriers that were only gradually dismantled over the following decade.

a big market share. Favara and Imbs (2015) show that the liberalization of the branching regime had a big impact on the growth rate of mortgage loans and, on housing prices. Our point is complementary to these two papers: we suggest that the political pressure to remove geographical restrictions on bank branching is likely to have increased as capital inflows into the U.S. financial system made geographical diversification more attractive for integrated banks.

Note that this is also what our theory would predict. As we show in a model extension in Appendix C, the marginal benefit for banks to increase their geographical diversification increases as capital inflows lower bank's refinancing rates and allow them to increase their leverage. Hence, we could expect the extensive margin of lending—new lending in new local markets—to be an important component of the credit expansion of the late 1990s and early 2000s.

Another dimension of possibly endogenous geographical diversification might be the increasing use of securitization in the years before the crisis. Very much as bank branching, securitization of mortgage loans may help banks reduce their exposure to locally concentrated risks, thus allowing the bank to lower its notional portfolio risk, absorb additional debt and increase lending.

In Table 8 we shed some light on these questions. Column 1 presents a cross-sectional regression of the post-1995 state-level average value of the interstate branching indicator IB_t^s (denoted by $\overline{IB}^s = \sum_{t>1995} IB_t^s / T$) on $OPEN^s \times CAPFLOW_t$. The coefficient is highly significant. This lends further support to the point made by Rice and Strahan (2010) that states that were more open already were more likely to dismantle remaining branching restrictions. In the

second column, we provide a state-level panel regression of the time-varying index IB_t^s on $OPEN^s \times CAPFLOW_t$ and our usual set of controls. This coefficient is also positive and highly significant, in line with our hypothesis. Capital inflows made cheap funding available to geographically diversified banks that in turn increased political pressure to liberalize interstate branching—in particular in states where they already had a strong foothold.

The last three columns of Table 8 provide some insight into how capital inflow shocks for banks affected securitization at the county-level. We obtain data on the number of loans sold in the county, broken down by banks and non-banks from the data set by Favara and Imbs (2015). We then regress these variables on $CAPFLOW_t^c$ and our usual set of county-level controls. Interestingly, capital inflow shocks lead banks to securitize fewer loans, while there is again no impact in the number of loans sold by non-banks. One possible interpretation of these findings is that geographically diversified banks securitized fewer loans in markets which they could enter relatively early and in which they had therefore established a strong foothold by the time of the savings glut.

5.4. Further independent bank-level evidence

We provide further independent evidence on the transmission mechanism based on quarterly Call Reports for the period 1997-2005. The call reports have rich bank-level information that is not contained in the HMDA. Specifically, besides aggregate mortgage lending, they contain data about bank's aggregate deposits, and equity as well as their lending and refinancing rates.

We construct interest rates from the item “interest and fee income on loans secured by real estate” (RIAD4011) and divide them by the stock of outstanding mortgage loans. Since the interest and fee income series is very noisy, we take four-quarter moving averages and divide them by the moving average of the banks real estate lending over the same period. Refinancing rates are computed in the same way using the item “interest on deposits” (RIAD4170) and bank deposits. For the analysis in this section, we proxy a bank’s degree of geographical diversification using an indicator of whether a bank is part of multi-bank holding company (MBHC).¹⁸

With these bank-level data, we then run local linear projections of the of the form

$$x_{t+h}^b = \alpha_h \text{MBHC}_t^b \times \text{CAPFLOW}_t + \text{CONTROLS} + \delta_b + \tau_t + \varepsilon_{t+1}^b$$

where x_{t+h}^b is the h-period cumulated change in one of the bank-level outcome variables discussed above (mortgage lending, interest rates, deposits), MBHC_t^b is a dummy indicating whether bank b is part of a MBHC, and δ_b and τ_t are bank and time effects. We then collect the estimated coefficients α_h for different forecasting horizons h to describe the dynamic responses of these variables to capital inflows. This local-linear-projection method was first suggested by

18. The downside of using the call reports is that the geographical information about *where* banks lend is rather coarse. While the information is precise at the state-level until the mid-1990s, the Riegle-Neal Interstate Banking and Branching Efficiency Act allowed banks to transform their local subsidiaries into branches. As a consequence, lending of these subsidiaries is recorded at the BHC-level and information about the the states where a bank is active is gradually lost.

Jorda (2005) and has the advantage that the response of x to fluctuations in CAPFLOW can be potentially non-linear in the time horizon h .

The findings, summarized in Figure 4, provide additional evidence on our mechanism: capital inflows impacted house prices via the lending policies of integrated banks. An increase in aggregate capital inflows leads to a decline in the bank's refinancing rate while deposits and lending increase. This finding of a simultaneous increase in liabilities and a decline in the refinancing rate provides additional independent evidence that capital inflows must have acted as a supply shock to geographically diversified banks. Equally consistent with the supply shock interpretation, we see that the increase in funding and the concomitant increase in lending gets partially passed on to mortgage borrowers via reduced mortgage rates.

6. Conclusion

In this paper, we studied the interaction between global imbalances in capital flows and interstate banking deregulation in the United States. We have argued that the tide of capital inflows that started to hit the United States from the middle of the 1990s onwards had a bigger impact on house prices in states that opened up their banking markets earlier during the 1980s and that therefore had a stronger presence of integrated banks, operating in several states, by the mid-1990s. Since aggregate inflows are exogenous with respect to local (state- and county-level) outcomes and since we use *ex ante* measures of financial integration—the number of years elapsed until 1995 since a state liberalized its local banking market to access from other states—this result

allows us to establish a causal link between aggregate capital inflows and state-level housing prices. Our results are robust to controlling for other common factors that could have affected house prices differently in different states such as low monetary interest rates or other indicators of credit availability or credit demand. We buttress this causal interpretation at several levels: using data on individual banks' lending in each U.S. county we show that geographically more diversified banks increased their lending in response to capital inflows and lowered their interest rates. Then, aggregating up individual bank's decisions for all counties we show that counties which—due to the specific deregulation histories of the states they are part of—had a stronger presence of geographically diversified banks, were more exposed to capital inflow shocks, saw larger increase in mortgage lending and higher house price growth.

Our findings contribute to a literature that has interpreted global imbalances as a the reflection of a global demand for safe assets—as a savings glut. We argue that this global demand for U.S. safe assets constituted a funding shock to the U.S. banking system. As opposed to purely local banks that operate only in one locality, integrated banks held a geographically diversified portfolio of mortgages. Since the aggregate U.S. housing market was considered safe at the time, this portfolio allowed them to tap the global demand for safe assets by refinancing themselves at low rates and by providing the international capital market with private-label safe assets in the form of mortgage-backed securities while increasing leverage at the same time.

Our results provide empirical support to the view that increased global demand for safe assets can actually lead to an increase in the prices of risky assets (Caballero and Krishnamurthy (2009)). Our finding suggests that the run-up in U.S. housing valuations occurred because regionally diversified financial intermediaries were perceived as safer than local banks. This allowed them to increase leverage and to invest into local mortgages, thus driving up housing markets. Hence, *intra*-national banking liberalization within the United States had a long shadow. It effectively increased the ability of the U.S. financial system to produce safe assets when a huge demand for such assets arose more than a decade later, with the emergence of China on the world economic stage.

TABLE 1. Summary Statistics

Variable	All states (1)				Early deregulators (by 1985) (2)				Late deregulators (after 1988) (3)			
	Mean	Median	Std. Dev.		Mean	Median	Std. Dev.		Mean	Median	Std. Dev.	
Years since deregulation	7.532	8.00	3.36		11.071	10.000	1.940		3.385	4.000	2.181	
House Prices Growth 1997-2012	0.041	0.040	0.055		0.040	0.040	0.068		0.040	0.038	0.035	
House price-income ratio 1997-2012	1.351	1.356	0.104		1.314	1.323	0.129		1.364	1.342	0.067	
$\Delta \ln(\text{State Personal Income})$ 1997-2012	0.009	0.009	0.012		0.009	0.009	0.012		0.010	0.010	0.012	
$\Delta \ln(\text{Population})$ 1997-2012	0.002	0.002	0.001		0.003	0.002	0.001		0.002	0.002	0.001	
Pre-sample trends												
(averages (means) over 1985 to 1996)												
House Prices Growth	0.049	0.049	0.039		0.050	0.050	0.053		0.050	0.049	0.022	
$\Delta \ln(\text{Tradable real GDP per capita})$	0.042	0.031	0.068		0.030	0.021	0.052		0.056	0.040	0.079	
$\Delta \ln(\text{Non-tradable real GDP per capita})$	0.043	0.044	0.035		0.042	0.039	0.042		0.046	0.044	0.029	
Real per capita GDP growth volatility	0.023	0.022	0.008		0.023	0.022	0.004		0.023	0.021	0.010	
$\Delta \ln(\text{Total employment})$	0.021	0.020	0.015		0.022	0.021	0.018		0.020	0.021	0.011	
Unemployment rate	5.959	5.873	1.152		5.560	5.661	1.164		5.493	5.615	0.977	
$\Delta \ln(\text{Tradable resid. wages})$	0.043	0.044	0.026		0.040	0.035	0.024		0.048	0.048	0.024	
$\Delta \ln(\text{Non-tradable resid. wages})$	0.060	0.061	0.026		0.065	0.066	0.033		0.060	0.060	0.018	
$\Delta \ln(\text{Total real estate lending})$	0.023	0.022	0.040		0.024	0.024	0.056		0.023	0.022	0.027	
Exports to GDP ratio	2.222	2.253	0.705		2.305	2.021	0.841		2.263	2.426	0.796	
Import exposure	0.056	0.053	0.032		0.068	0.054	0.048		0.055	0.057	0.024	
Current account to GDP (1997-2012)												
Current account to banking equity ratio (1997-2012)	0.038	0.038	0.013									
Current account to banking equity ratio (1997-2012)	0.152	0.164	0.055									

The upper panel of the table shows means, medians and standard deviations of main variables and baseline controls. For each variable and group of state (early or late deregulators) we compute the cross-state mean and median of the (within-state, time-series) mean as well as the average of the state-level standard deviations. The panel in the middle provides descriptives for the pre-sample period (1985-1996). The lower panel shows descriptives for the aggregate variables. Note that the variables *Years since deregulation*, *Import exposure*, *Exports to GDP ratio* and *Real per capita GDP growth volatility* do not vary over time, so that mean, median and standard deviations are just calculated across states.

TABLE 2. House Prices, Financial Openness and Capital Inflows — state-level results

Dependent variable is house price growth , $HPGROWTH_t^S$					
	(1)	(2)	(3)	(4)	(5)
$OPEN^S \times CAPFLOW_t$	0.030 (2.76)	0.030 (3.19)	0.044 (3.47)	0.032 (2.83)	0.026 (1.42)
BASELINE		YES	YES	YES	YES
$OPEN^S \times CREDIT_t$			YES	YES	
$PRE97^S \times CAPFLOW_t$				YES	
$CREDIT_t \times State$					YES
$PRE97^S \times Time$					YES
Observations	2820	2820	2820	2820	2820
R^2	0.60	0.70	0.71	0.72	0.89

The Table shows the results from the panel regression

$$HPGROWTH_t^S = \alpha \times OPEN^S \times CAPFLOW_t + \text{CONTROLS} + \tau_t + \delta^S + \varepsilon_t^S,$$

where $OPEN^S = 1995 - \text{Year of Banking Deregulation}$ and $CAPFLOW$ are capital inflows, measured as the negative current account balance over lagged total U.S. banking equity ($CAPFLOW_t = -CA_t / BankEquity_{t-1}$). The dependent variable, $HPGROWTH_t^S$, is the four-quarter ahead change in house prices. The baseline set of controls includes the growth rates of per capita state personal income, and population growth, the house price income ratio and the inter-state-branching indicator by Rice and Strahan (2010). The vector $CREDIT_t$ includes variables capturing aggregate credit demand and supply: the real 10-year treasury constant maturity rate, the real 3-months T-bill rate, the Baa-Aaa (default) spread, the survey of lending attitudes of senior loan officers and the Michigan survey of consumer house buying conditions index.

The vector $PRE97^S$ includes 1985-1996 means of the following state-level variables: house price growth, the growth rates of tradable and non-tradable GDP and wages, volatility of GDP growth, employment growth, the unemployment rate, real estate lending growth, exports relative to GDP and a measure of import exposure (Autor et al. (2013)) aggregated to the state-level.

Data are quarterly, sample period is 1997-2012. Sample includes 47 US states (Alaska, Hawaii, Delaware as well as District of Columbia are left out). OLS estimates, all regressions include time- and state-fixed effects. Standard errors are clustered by state and time, t-statistics appear in parentheses.

TABLE 3. Robustness across sub-samples; Placebo treatments

Dependent variable is house price growth, $HPGROWTH_t^s$							
1997-2012 on subset of states				47 states on subperiods		$open_{placebo}^s$	
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
w/o sand-states	w/o MT, KS, ND, NE, ID	w/o ME, SD	w/o CA, NY	1997-2005	2006-2012	1985-1996	1997-2012
$OPEN^k \times CAPFLOW_t$	0.033 (3.19)	0.041 (3.51)	0.030 (2.48)	0.032 (1.25)	0.026 (1.55)	0.002 (0.23)	simulated coefficients mean std 0.00004 0.013
CONTROLS	YES	YES	YES	YES	YES	YES	fraction of significant coefficients 0.05
Observations	2580	2700	2700	1692	1128	2068	
R^2	0.75	0.72	0.73	0.55	0.80	0.71	

The Table shows the results from the panel regression

$$HPGROWTH_t^s = \alpha \times OPEN^k \times CAPFLOW_t + \text{CONTROLS} + \tau_t + \delta^s + \varepsilon_t^s,$$

on the sub-samples indicated in the column heading. Again, $OPEN^k = 1995 - \text{Year of Banking Deregulation}$ and $CAPFLOW$ are capital inflows, measured as the negative current account balance over lagged total U.S. banking equity ($CAPFLOW_t = -CA_t / BankEquity_{t-1}$). The dependent variable, $HPGROWTH_t^s$, is the four-quarter ahead growth in house prices. **CONTROLS** comprises the full set of controls from Table 2: **BASLINE**, $CREDIT \times OPEN^s$, and $PRES^s \times CAPFLOW_t$. The vector **BASLINE** includes the growth rates of per capita state personal income, and population growth, the house price income ratio and the inter-state-branching indicator by Rice and Strahan (2010). The vector **CREDIT**, includes variables capturing aggregate credit demand and supply: the real treasury constant maturity rate, the real 3-months T-bill rate, the Baa-Aaa (default) spread, the survey of lending attitudes of senior loan officers and the Michigan survey of consumer house buying conditions index. The vector **PRES^s** includes 1985-1996 means of the following state-level variables: house price growth, the growth rates of tradable and non-tradable GDP and wages, volatility of GDP growth, employment growth, the unemployment rate, real estate lending growth, exports relative to GDP and a measure of import exposure (Autor et al. (2013)) aggregated to the state-level. Data are quarterly, sample period in columns (1)-(4) and (8) is 1997-2012. OLS estimates, all regressions include time- and state-fixed effects. Standard errors are clustered by state and year, t-statistics appear in parentheses.

The last column reports results from the Monte Carlo simulations for $OPEN^s$. We take 10000 random draws from the empirical distribution of $OPEN^s$. The reported numbers are the average simulated coefficient and the fraction of cases where estimated coefficients are individually significant at the 5% level.

TABLE 4. State liberalization and bank characteristics

	bank geographical diversification			dependent variable			bank balance sheet characteristics	
	interstate asset ratio (1)	interstate mortgage ratio (2)	number of states where bank operates (3)	1/HFI (4)	equity ratio (5)	wholesale funding (6)	retail deposits (7)	
OPEN ^s	0.04 (4.51)	0.03 (4.21)	0.13 (1.97)	0.06 (3.68)	-0.003 (-3.84)	0.003 (1.65)	-0.009 (-2.47)	
Observations	47	47	47	47	47	47	47	
R ²	0.30	0.27	0.06	0.21	0.23	0.04	0.10	

The table presents cross-sectional state-level OLS regressions of the characteristics X^s of the state-level banking system on baseline measure of openness, OPEN, the number of years elapsed until 1995 since state-level liberalization:

$$X^s = \alpha_0 + \alpha_1 \times \text{OPEN}^s$$

X^s is the mean over 1986 to 1993 of $X_t^s = \sum_{b=1}^B \lambda_t^b \times \omega_{t-1}^{b,s}$ where $\omega_{t-1}^{b,s}$ is the share of banking assets (or, in column (2), mortgage assets) held in state s by bank b and λ_t^b is one of the following bank characteristics: in columns (1) and (2) a dummy for whether the bank is part of a bank holding company that is active in other states (so that X_t^s is the share of assets or mortgages owned by interstate banks), in column (3) the number of states in which a bank is operating, in column (4) $1/HFI^b$ where HFI^b is a measure of the Herfindahl index measuring the concentration of a bank's mortgage portfolio across states. In column (5) the equity ratio, in column (6) the ratio of wholesale funding to total assets and in column (7) the ratio of retail deposits to total assets. Numbers in parentheses are t-statistics.

TABLE 5. Bank-level regressions, 1997-2005

	dep. variable: β_t^b			dep. variable: ∇_t^b		
	σ_n at level			σ_n at level		
	county	MSA	State	county	MSA	State
	(1)	(2)	(3)	(4)	(5)	(6)
$RTC_{t-1} \times CAPFLOW_t$	4.360 (1.267)	31.362 (2.807)	33.645 (2.453)	0.122 (2.224)	0.488 (3.030)	1.637 (3.385)
log total assets	0.008 (0.020)	-0.004 (-0.009)	0.093 (0.171)			
Bank Asset Growth	0.003 (0.972)	0.003 (0.391)	0.003 (1.102)			
RTC_{t-1}	-4.055 (-1.312)	-28.798 (-2.777)	-35.645 (-2.213)	-0.115 (-1.518)	-0.456 (-1.747)	-1.535 (-1.624)
Observations	31,732	31,732	31,732	31,732	31,732	31,732
R ²	0.076	0.082	0.076	0.001	0.001	0.001

The first three columns of the table present bank-level regressions of the bank lending supply shocks β_t on bank characteristics and aggregate factors of the form

$$\beta_t^b = \alpha \times RTC_{t-1} \times CAPFLOW_t + \text{CONTROLS} + \delta^b + \tau_t + \nabla_t^b$$

where $CAPFLOW$ are capital inflows, measured as the negative current account balance over lagged total U.S. banking equity ($CAPFLOW_t = -CA_t / BankEquity_{t-1}$). **CONTROLS** includes the logarithm of bank total assets, aggregate lending growth and the interaction $RTC_{t-1} \times \text{CREDIT}_t$ with **CREDIT** the vector of aggregate credit supply factors of Favilukis et al. (2012). Bank's risk taking capacity RTC_{t-1} is computed according to equation (10), with σ_{t-1} calculated according to equation (12) at the county (column (1)), MSA (column (2)) and state levels (column (3)) respectively, $E(r^l - r^d) = r^d = 0.01$ and $\Phi = 2.32$.

The right panel (columns (4)–(6)) presents the regressions of the residual ∇_t^b of the respective regression in column (1)–(3) (re-run with $RTC_{t-1} \times CAPFLOW_t$ dropped) on $RTC_{t-1}^b \times CAPFLOW_t$, RTC_{t-1} and bank and time fixed effects. All regressions are WLS using lagged U.S.-wide market share of banks as weights. Standard errors are clustered by bank and year. Numbers in parentheses are t-statistics.

TABLE 6. County-level regressions of house prices on $CAPFLOW_t^c$

	(1)	(2)	(3)	(4) state-time	(5) border-MSA	(6) GMM
$CAPFLOW_t^c$	4.908 (8.355)	4.372 (5.205)	2.086 (3.041)	2.302 (1.150)	1.294 (1.910)	0.896 (4.483)
$CAPFLOW_t^c \times ELASTICITY^c$			0.231 (1.035)	-0.164 (-2.130)	-0.125 (-0.569)	
$IB_{t-1}^{s(c)}$		-0.009 (-1.262)	0.012 (1.818)		0.011 (2.370)	0.006 (10.185)
$IB_{t-1}^{s(c)} \times ELASTICITY^c$			-0.007 (-3.998)		-0.005 (-1.940)	-0.003 (-12.273)
$HPGROWTH_{t-1}^c$			0.418 (10.675)	0.321 (19.447)	0.449 (13.360)	0.592 (30.420)
income growth		0.351 (5.961)	0.320 (8.556)	0.321 (8.065)	0.079 (0.695)	0.024 (0.976)
income growth lagged		0.373 (12.115)	0.083 (2.636)	0.172 (13.131)	0.052 (0.447)	0.098 (6.461)
population growth		0.820 (3.768)	1.264 (4.490)	0.874 (2.149)	0.194 (0.744)	0.245 (5.582)
population growth lagged		2.136 (2.770)	0.841 (1.894)	0.905 (2.939)	1.293 (1.965)	-0.008 (-0.203)
market concentration		0.029 (2.467)	0.033 (2.547)	0.026 (1.345)	0.005 (0.919)	0.003 (2.347)
market concentration lagged		0.007 (1.455)	0.003 (0.792)	0.010 (1.380)	0.008 (1.227)	0.002 (1.970)
credit demand (γ_t^c)		0.015 (1.369)	0.012 (1.068)	0.020 (1.773)	-0.006 (-0.392)	
Observations	11,852	9,872	8,612	8,612	2,440	1,042
R ²	0.404	0.539	0.629	0.818	0.571	Sargan-test 426.27 (p=0.00)

The table shows results of house prices regressions of the form

$$HPGROWTH_t^c = \alpha_0 \times CAPFLOW_t^c + \alpha_1 \times ELASTICITY^c \times CAPFLOW_t^c + \mathbf{CONTROLS} + \delta_c + \tau_t + \varepsilon_t^c$$

for urban (MSA) counties for the period 1997-2005. $ELASTICITY$ is the housing supply elasticity by Saiz (2010) and $IB^{s(c)}$ the state-level indicator of interstate branching deregulation compiled by Rice and Strahan (2010). All regressions are weighted with the (lagged) share of a county in U.S. aggregate mortgage lending and have county and time effects. The regression in column (4) has state-time fixed effects. The regression in column (5) constrains the sample to a subset of counties in MSAs that straddle state borders. See Table D.6 for a list of these MSAs. Standard errors for regressions are clustered by year of liberalization ($OPEN^{s(c)}$) of the state the county is located in. Numbers in parentheses (columns (1)–(5)) are t-statistics. In column (6) we report results of a system GMM estimation based on Blundell and Bond (1998). Numbers in parentheses in column (6) are z-scores.

The county-level capital inflow shock $CAPFLOW_t^c$ is constructed from the fitted values of the residual regressions in column (5) of Table 5, i.e. using MSA-level diversification.

TABLE 7. County-level regressions of lending growth on $CAPFLOW_t^c$ and IV results

	Dep variable: mortgage lending growth			Dep variable: HPGROWTH $_t^c$	
	(1) all institutions	(2) Banks	(3) Non-Banks	(4) IV	(5) IV
Lending growth by banks fitted				0.448 (3.600) [F=17.05] [p<0.001]	0.406 (2.236) [F=4.15] [p<0.05]
Lending growth by banks $\times$ ELASTICITY fitted					0.032 (0.484) [F=38.93] [p<0.01]
$CAPFLOW_t^c$	4.203 (1.226)	5.676 (4.034)	0.506 (0.364)		
$IB_{t-1}^{s(c)}$	-0.003 (-0.256)	-0.014 (-2.769)	0.013 (4.070)	0.003 (0.884)	0.017 (2.093)
$IB_{t-1}^{s(c)} \times ELASTICITY$					-0.006 (-1.765)
$HPGROWTH_{t-1}^c$				0.457 (11.961)	0.438 (14.714)
income growth	-0.045 (-0.398)	0.113 (3.004)	0.089 (1.546)	0.271 (5.111)	0.271 (4.734)
income growth lagged	0.330 (2.711)	0.004 (0.049)	0.017 (0.126)	0.057 (0.998)	0.069 (2.054)
population growth	2.088 (3.035)	-0.397 (-1.604)	1.633 (2.060)	1.489 (4.041)	1.382 (3.776)
population growth lagged	1.566 (1.894)	0.007 (0.017)	0.088 (0.134)	0.747 (3.172)	0.825 (3.382)
market concentration	0.028 (1.306)	-0.006 (-0.522)	0.021 (1.077)	0.037 (3.205)	0.034 (4.413)
market concentration lagged	-0.080 (-3.199)	-0.021 (-1.402)	-0.026 (-2.767)	0.011 (1.231)	0.010 (1.146)
credit demand (γ_t^c)	0.413 (7.402)	0.046 (1.571)	0.141 (6.057)	-0.009 (-0.659)	-0.006 (-0.350)
Observations	9,872	9,872	9,872	9,872	8,612
R ²	0.933	0.774	0.883	0.438	0.437

Columns (1)–(3) of the table show results of regressions of the form

$$\frac{\Delta L_t^c}{L_{t-1}^c} = \alpha_0 \times CAPFLOW_t^c + \text{CONTROLS} + \delta_c + \tau_t + \epsilon_t^c$$

where $\Delta L_t^c / L_{t-1}^c$ is mortgage lending growth in county c . The first column provides the results for total mortgage lending growth (by all institutions active in the county), column (2) for commercial banks and column (3) for non-banks. $IB_{t-1}^{s(c)}$ is the state-level indicator of interstate branching deregulation by Rice and Strahan (2010). The county-level capital inflow shock $CAPFLOW_t^c$ is constructed from the fitted values of the residual regressions in column (5) of Table 5, i.e. using MSA-level diversification.

Column (4) and (5) instrument lending growth by commercial banks with $CAPFLOW_t^c$ in house price regressions of the form

$$HPGROWTH_t^c = \alpha_0 \times \widetilde{\frac{\Delta L_t^c}{L_{t-1}^c}} + \alpha_1 \times ELASTICITY^c \times \widetilde{\frac{\Delta L_t^c}{L_{t-1}^c}} + \text{CONTROLS} + \delta_c + \tau_t + \epsilon_t^c$$

where the tilde denotes the fitted value from the first stage. All regressions are weighted with the (lagged) share of a county in U.S. aggregate mortgage lending. Standard errors are clustered by year of liberalization ($OPEN^{s(c)}$) of the state the county is located in. Numbers in round parentheses are t-statistics. The numbers in rectangular brackets in column (4) denote the first-stage F-statistics and the associated p-value. In column (5), rectangular brackets show tests for weak identification based on the conditional F-statistics by Sanderson and Windmeijer (2016). The associated p-values are bootstrapped. The sample includes all urban counties (i.e. counties that are part of an MSA) for the period 1997–2005.

TABLE 8. Extensive margins of diversification and capital inflows

	Interstate branching		Securitization		
	(1) Cross- sectional	(2) State- panel	(3) All banks	(4) Out-of- state banks	(5) non- banks
OPEN ^s	0.202 (5.383)				
OPEN ^s × CAPFLOW _t		1.181 (3.288)			
CAPFLOW _t ^c			-10.402 (-2.874)	-10.829 (-2.767)	-2.094 (-0.354)
CONTROLS		YES	YES	YES	YES
Observations	47	2,068	8,570	8,531	8,483
R ²	0.392	0.826	0.316	0.351	0.479

The first two columns illustrate the impact of capital inflows on interstate branching. The first column reports a cross-sectional state-level regression of the form $\overline{IB}^s = \alpha_0 + \alpha_1 \text{OPEN}^s + \varepsilon_s$ where $\overline{IB}^s = \frac{1}{T} \sum_{t=1}^T IB_t^s$ is the average post-1997 degree of interstate-branching deregulation.

The second column reports a state-level panel regression $IB_t^s = \alpha_0 + \alpha_1 \text{OPEN}^s \times \text{CAPFLOW}_t + \text{CONTROLS} + \delta_s + \tau_t + \varepsilon_t^s$.

Data are quarterly, sample period is 1995-2005. Sample includes 47 US states (Alaska, Hawaii, Delaware as well as District of Columbia are left out). OLS estimates.

Columns (3)–(5) report regressions of the (growth in) the number of loans sold for securitization on the county-level capital inflow shock CAPFLOW_t^c constructed from the fitted values of the residual regressions in column (5) of Table 5, i.e. using MSA-level diversification and the full set of controls also used in Tables 6 and 7: interstate branching, current and lagged values of population and income growth and the index of market concentration as well as the estimate of the county-level loan demand shock, γ_t^c . Column (3) has the growth in the total number of loans sold by banks as dependent variable. Column (4) the growth in the number of loans sold by out-of-state banks and column (5) the number of loans sold by non-banks. Data are annual, 1997-2005. Regressions are weighted with the lagged share of the county in aggregate mortgage lending as weight. Standard errors of county-level regressions are clustered by year of liberalization of the state the county is located in (OPEN^s). T-statistics appear in parentheses.

FIGURE 1. Capital Inflows and House Price Growth, 1985-2012

NOTES: The figure plots the ratio of the U.S. current account to total banking equity of U.S. commercial banks (solid line) against cross-state averages of four-quarter-ahead house price growth. The dot-dashed line is the average house price growth for states that deregulated their interstate banking early (by 1985). The long-dashed line represents average house price growth for states that deregulated late (after 1988). The vertical line indicates the 4th quarter of 1996.

FIGURE 2. House Price Growth and Banking Deregulation—cross-sectional evidence

NOTES: The figure plots the averages of $HPGROWTH_{subperiod}^s$ for each state s for three sub-periods, pre-97, 1997-2005 and 2006-2012, against ex-ante openness $OPEN^s$, which is defined as $OPEN^s = 1995 - Year\ of\ Banking\ Deregulation$. $HPGROWTH_t^s$ is the four-quarter ahead change in house prices. Cross-sectional regressions of state averages of house price changes on the baselines openness measure, $OPEN^s$ $HPGROWTH_t^s = \beta OPEN^s + constant + \varepsilon_s$, yield the following coefficient estimates (t-statistics) and R^2 for the subperiods : pre-97: 0.0001 (0.26) $R^2=0.00$; 1997-2005: 0.0024 (2.64) $R^2=0.11$, 2006-2012: -0.0034 (-2.63) $R^2=0.11$.

FIGURE 3. House Price Growth and Capital Inflows—time series evidence by state

NOTES: The figure plots the estimate of the coefficient α^s for each state s from the state-by-state time-series regression $\text{HPGROWTH}_t^s = \alpha^s \times \text{CAPFLOW}_t + \delta_s + \text{CONTROLS} + \varepsilon_{s,t}$ against ex ante openness OPEN^s , which is defined as $\text{OPEN}^s = 1995 - \text{Year of Banking Deregulation}$ and CAPFLOW are capital inflows, measured as the negative current account balance over lagged total U.S. banking equity ($\text{CAPFLOW}_t = -\text{CA}_t / \text{BankEquity}_{t-1}$). The dependent variable, HPGROWTH_t^s , is the four-quarter ahead growth in house prices. The solid line is based on the cross-sectional regression $\alpha^s = a + b \times \text{OPEN}^s + u_s$ $b = 0.06$ (4.43) $R^2 = 0.30$. T-statistic is in parentheses.

FIGURE 4. Dynamic Responses in Lending Growth and Interest Rate

NOTES: The figure plots dynamic responses of bank lending and deposits (upper panel), interest and deposit rates (lower panel) to the movements in U.S. capital inflows as a function of bank's geographical diversification. Geographical diversification is measured by the multi-bank holding company dummy, $MBHC_t^b$. The dynamic responses represent the sequence of coefficients α_h in regressions of the form

$$x_{t+h}^b = \alpha_h MBHC_t^b \times CAPFLOW_t + TotalAssets_t^b + \delta^{BHC} + \tau_t + \varepsilon_t^b$$

where x_{t+h}^b is the h -period ahead interest/deposit rate of bank b or the cumulated lending/deposit growth of this bank between period t and $t+h$. δ^{BHC} denotes MBHC-fixed effects (or the bank-fixed effect for banks that are not part of MBHC). Standard errors are clustered by MBHC and time.

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Online Appendix for

Mathias Hoffmann and Iryna Stewen

**“HOLES IN THE DIKE THE GLOBAL SAVINGS GLUT, U.S. HOUSE PRICES
AND THE LONG SHADOW OF BANKING DEREGULATION”**

Appendix A: Appendix: Data

Our empirical analysis uses a panel of variables for the 47 contiguous U.S. states excluding Delaware for the period 1997-2012.

House price growth, $HPGROWTH_t^s$, is our main dependent variable and is calculated as the four-quarter ahead growth rate in house prices.

House prices. We use quarterly data on land and property values provided by the U.S., Lincoln Institute of Land Policy. The data are based on the adjusted Federal funding housing agency (FHFA) indexes estimated for 50 U.S. states and the District of Columbia of the prices and quantities of residential housing and its two components, land and structures. Calculated using the method by Davis and Heathcote (2007), these data are uniquely suited for our analysis since they are corrected for interstate differences in the quality of residential land and housing and are thus comparable across states.

State personal income is quarterly per capita personal income by state provided by the Bureau of Economic Analysis (BEA). Growth rates are calculated as the first differences of the natural log of level values.

House price-income ratio, HPY_t^s , is the logarithm of the ratio of house prices to personal income.

Population is quarterly population by state provided by the Bureau of Economic Analysis (BEA). Growth rates are calculated as the first differences of the natural log of level values.

State-level financial openness. Our main measure, $OPEN^s$, indicates how many years had passed by 1995 since interstate banking deregulation took

place, i.e.

$$\text{OPEN}^s = 1995 - \text{Year of Banking Deregulation}$$

Deregulation dates are from Kroszner and Strahan (1999). In the few cases where states maintained internal barriers to branching even after allowing in banks from outside the state, we use the minimum number of years that have passed until 1995 since either interstate banking or intra-state branching deregulation. For comparison, we also compute several de facto measures of banking market integration based on data from the Call reports that we describe in more detail below.

Interstate Branching (IB_i^s). The index is constructed using information provided in the Table 1 of Rice and Strahan (2010) on the effective date of interstate branching regulation changes, and each of the following four provisions: the minimum age of the institution for acquisition, allowance of de novo interstate branching, allowance of interstate branching by acquisition of a single branch or portions of an institution, and statewide deposit cap on branch acquisitions. The index is set to zero for states that impose all four restrictions to out-of state-entry. Abolishment of each of the restriction adds one quarter to the index. While the original index by Rice and Strahan (2010) ranges from 0 to 4, we re-normalize it to range from 0 (no integration) to 1 (full integration).

Capital Inflows (CAPFLOW_i). Our first and principal measure is the (negative) current account balance over lagged total U.S. banking equity. The current account balance is from the BEA, U.S. International Transactions Accounts Data, quarterly and seasonally adjusted. U.S. banking equity data

is quarterly Total Equity Capital for Commercial Banks in United States [USTEQC] from the Federal Financial Institutions Examination Council (US), retrieved from FRED, Federal Reserve Bank of St. Louis. Alternatively we measure ($CAPFLOW_t$) by net U.S. incurrence of liabilities excluding financial derivatives (net increase in liabilities / financial inflow (+)) and net U.S. acquisition of financial assets excluding financial derivatives (net increase in assets / financial outflow (+)) and define them respectively as $INFLOW_t$ and $OUTFLOW_t$. These two time-series are provided by BEA international transaction statistics and are also expressed relative to U.S. banking equity. Another alternative measures is *net foreign holdings of total assets* defined as foreign-owned assets in the United States minus U.S.-owned assets abroad and *net foreign holdings of total securities* is defined as foreign-owned U.S. government securities plus U.S. Treasury securities plus U.S. securities other than Treasury securities minus U.S.-owned foreign securities. The last two measures of capital inflows are quarterly and provided by the BEA, U.S. International Transactions Accounts Data. They are expressed relative to nominal GDP at current prices.

Indicators of monetary policy and credit availability ($CREDIT_t$): The *short-term real interest rate*, is constructed as U.S. (effective) Federal Funds minus U.S.-wide inflation. Data on Federal Funds are from the Board of Governors of the Federal Reserve System, Historical Data. The data is monthly and to compute quarterly data we average it over 3 months. U.S. inflation is computed using quarterly data on Personal Consumption Expenditures from the BEA. Our measure of *monetary policy looseness* is constructed as the

deviation of the monetary policy rate from the interest rate implied by a Taylor rule where the monetary policy rate is the U.S. (effective) Federal Funds rate from the Board of Governors of the Federal Reserve System. The Taylor rule we use is: $0.02 + 1.5(\pi - 0.02) + 0.5 \times \text{output gap}$, where π is U.S.-wide inflation and the output gap is measured by detrending an index for real GDP (constructed using the cumulation of official quarterly real GDP growth rates) with the HP-filter. *Real long-term interest rates* are measured as the 10-year constant maturity Treasury bond rate minus expectations of the average annual rate of CPI inflation over the next 10 years from the Survey of Professional Forecasters (only available from 1992), in percent per annum. Finally, we use two measures from surveys on credit conditions. The first is Net Percentage of *Domestic Banks Reporting Increased Willingness to Make Consumer Installment Loans* [DRIWCIL] from the Senior Loan Officer Opinion Survey on Bank Lending Practices of the Board of Governors of the Federal Reserve System (US), retrieved from FRED, Federal Reserve Bank of St. Louis. A positive value for this variable therefore indicates a loosening of credit conditions. The second measure, “*Good time to buy*”, the University of Michigan Survey of Consumers that reflects consumers’ opinion towards the house buying conditions. *Financial Distress* is measured as Corporate Bond Yield Spread between AAA- and BAA-rated corporate bonds.

Pre-97 state characteristics (PRE97^s): GDP of tradable sector is measured by manufacturing sector gdp at state level from the BEA regional accounts. Data is quarterly.

GDP of non-tradable is measured by retail trade sector GDP at state level from the BEA regional accounts. Data is quarterly.

GDP growth volatility is standard deviation of real per capita state-level GDP growth from the BEA regional accounts. Data is quarterly.

Total employment is total full-time and part-time employment from the BEA regional accounts. Data is quarterly.

Unemployment rate is from the bureau of labor statistics.

Tradable residential wages are measured by private non-farm wages and salaries in manufacturing sector and are from the BEA regional accounts. Data is quarterly.

Non-tradable residential wages are measured by private no-farm wages and salaries in retail trade sector and are from the BEA regional accounts. Data is quarterly.

Total real estate lending is aggregated to the state level from the Call Reports using bank-level data on total loans secured by real estate (rcfd1410).

Exports to GDP ratio. State-level data on exports is calculated following using the 1993 Commodity Flow Survey (CFS) data on interstate shipments among the 48 contiguous US states.

Import exposure is constructed following Autor et al. (2013) but using 2-digit industry-level data at the state level as in Hoffmann and Khabibulina (2019).

Mean growth rates of pre-97 state variables are calculated as the first differences of the natural log of level values. We then take the mean over the period 1986Q1-1996Q4.

Bank-level data (call reports): Mortgage lending. Bank lending and interest rate on mortgage lending are computed using data from the Call Reports. The data are available for the period 1986-2010 on the quarterly basis. For each commercial bank the data provides us with information on identification

number (rssd 9001), total loans secured by real estate (rcfd1410), state of location (rssd9200), the BHC with which it is affiliated—if any—(rssd9348), and interest and fee income on loans secured by real estate. Mortgage rates are calculated as loans secured by real estate by the stock of outstanding mortgage loans. Both variables are smoothed before using 4-period moving average. Banks are divided into two groups depending if they are owned by a BHC that operates in several states –interstate or integrated banks– or belong to a in-state local bank –local banks. Real estate loans and interest and fee income for these two groups are then aggregated, each quarter, at the state level. We also aggregate these data to the BHC-level, for banks belonging to the MBHC.

Deposits. Bank's deposits and refinancing rates are computed analogously to mortgage lending and mortgage rates using data on bank total deposits (rcfd2200) and interest on deposits (riad4170).

Wholesale funding. Bank's total wholesale funding, which is the sum of wholesale deposits (brokered and foreign deposits, as well as large time deposits beyond \$100,00), fed funds, repo borrowing, and other borrowed money.

Retail deposits. Bank's retail deposits are calculated by subtracting wholesale deposits from total deposits.

Equity ratio. Bank's equity ratio is bank total equity (RCON3210) divided by bank total assets (RCON2710).

Appendix B: Appendix: the state-level regression in a reduced-form model of banks' credit supply

In this appendix, we show that our state-level regression

$$\text{HPGROWTH}_t^s = \alpha \text{OPEN}^s \times \text{CAPFLOW}_t + \tau_t + \delta_s + \varepsilon_t^s$$

(and its de facto counterparts in Table D.1), arise naturally from aggregating individual banks' credit supply decisions in an extended version of the the reduced-form model by Landier et al. (2017). Following Landier et al. (2017), we assume that lending supply in state drives local house prices growth with an elasticity of α , so that

$$\text{HPGROWTH}_t^s = \alpha \frac{\Delta L_t^s}{L_{t-1}^s} + v_t^s$$

where

$$L_t^s = \sum_{b=1}^B L_t^{b,s}$$

is state-level lending and $L_{n,t}^{b,s}$ is the lending of bank b in state s and B the number of banks. We assume that the growth rate of lending supply of bank b is given by

$$\frac{\Delta L_t^{b,s}}{L_{t-1}^{b,s}} = \gamma_t + \lambda_{t-1}^b \text{CAPFLOW}_t + \eta_t^{b,s}$$

where γ_t captures aggregate factors that affect all banks equally and $\eta_{b,t}^s$ is a bank and/or state-specific shock which is assumed to be uncorrelated with aggregate factors and capital inflows but may correlate with other banks' lending shocks as well as with the local house price shocks ν_t^s . The (potentially time-varying) parameter $\lambda_{t-1}^b \in \left[0, 1 \right]$ captures the exposure of bank b to aggregate capital inflows.

The key feature of this reduced-form specification—which it extends Landier et al. (2017)—is that banks differ in their exposure to the aggregate capital inflow factor CAPFLOW_t . This general reduced-form of the lending supply function is consistent in particular with the more articulated theoretical model that we describe in the paper and in Appendix C in which banks' lending supply is determined by a value-at-risk (VaR) constraint. In this model, more diversified banks will increase lending more in response to a given capital inflow, which is exactly what the reduced-form here is saying.

Based on this general setup, we can now write the house price change in state s as:

$$\text{HPGROWTH}_t^s = \alpha \frac{\Delta L_t^s}{L_{t-1}^s} + \varepsilon_t^s = \alpha \sum_{b=1}^N \frac{\Delta L_t^{b,s}}{L_{t-1}^{b,s}} \omega_{t-1}^{b,s} + \nu_t^s$$

where $\omega_{t-1}^{b,s} = L_{t-1}^{b,s} / L_{t-1}^s$ is the lending share of bank b in state s . Plugging in for $\Delta L_{n,t}^{b,s} / L_{t-1}^{b,s}$ from the bank lending supply equation, we get

$$\text{HPGROWTH}_t^s = \alpha \left(\sum_{b=1}^B \left(\lambda_{t-1}^b \text{CAPFLOW}_t + \eta_t^{b,s} + \gamma_t \right) \omega_{t-1}^{b,s} \right) + \nu_t^s$$

With this representation, it is now natural to define the financial openness of state s as the (market-share) weighted average of the geographical

diversification of banks:

$$\text{OPEN}_{t-1}^{\text{defacto},s} = \sum_{n=1}^N \lambda_{n,t-1} \omega_{n,t-1}^s$$

so that

$$\text{HPGROWTH}_t^s = \tau_t + \alpha \text{OPEN}_{t-1}^{\text{defacto},s} \times \text{CAPFLOW}_t + \varepsilon_t^s$$

where $\tau_t = \alpha \times \gamma_t$ and $\varepsilon_t^s = \alpha \left(\sum_{n=1}^N \eta_{n,t}^s \omega_{n,t-1}^s \right) + v_t^s$. In this setting, OPEN_{t-1}^s will be correlated with ε_t^s via the bank market shares $\omega_{t-1}^{b,s}$. To account for this source of endogeneity and for the possibility that capital inflows impact on banks diversification decisions, we use ex ante (pre-sample) averages of $\text{OPEN}_{t-1}^{\text{defacto},s}$ in our estimations in Table D.1. As we have shown in Table 4, our de iure measure of the number of years elapsed since liberalization, OPEN^s , is positively related to de facto openness / geographical diversification measures. Our de iure measure OPEN^s therefore effectively serves as an “instrument” for the geographical diversification of a state’s banks in our state-level regressions.

It is interesting to relate the model above to the findings of Landier et al. (2017), who emphasize the role of granularity in the bank-size distribution for the synchronization of house prices across states. In our setup, conditional on CAPFLOW_t , lending and house prices are perfectly correlated across states. Still, capital inflows differ in their impact on house prices across states if bank-size distributions are very different between states. This is because, unlike in Landier et al. (2017), the common factor CAPFLOW_t differs in its impact on banks, depending on the banks degree

of diversification. Our openness measure $\text{OPEN}_t^{\text{defacto},s}$ therefore captures two dimensions of granularity: the within-bank dimension that determines the banks' geographical diversification, λ_{t-1}^b , and the within-state dimension of the bank size distribution, reflected in banks' local market shares, $\omega_{t-1}^{b,s}$.

Appendix C: Appendix: A Value-at-risk model of bank lending and global imbalances

We use a variant of the model by Shin (2010) to describe the optimal leverage choice of individual banks. In a second step, we then consider equilibrium in a market for deposits in which there are two types of banks: integrated banks which are geographically diversified and local banks which are not. The maintained assumption in the model is that bank deposits are safe (due to deposit insurance or due to bailout expectations or because banks insolvency risk is sufficiently small). As in Caballero and Krishnamurthy (2009), the global demand for U.S. safe assets is exogenous and amounts to the “deposit” supply to the U.S. financial system. Global imbalances are modeled as an exogenous increase in this deposit supply.

Start from the balance sheet identity of the typical bank

$$A_t = E_t + D_t$$

where E_t denotes equity, D_t bank debt (deposits and wholesale debt) and A_t total bank assets (which we assume are just mortgage loans, so that $A_t = L_t$ throughout) at time t . The bank pays the safe refinancing (deposit) rate r_t^d and it generates a (risky) return r_{t+1}^l with standard deviation σ on its assets. Then

the bank's equity in the next period is

$$\begin{aligned}
E_{t+1} &= A_t(1 + r_{t+1}^l) - (1 + r_t^d) D_t \\
&= A_t(1 + r_{t+1}^l) - (1 + r_t^d) (A_t - E_t) \\
&= (r_{t+1}^l - r_t^d) A_t + (1 + r_t^d) E_t
\end{aligned}$$

The bank becomes insolvent whenever $E_{t+1} < 0$ which will happen whenever the return r_{t+1}^l falls below the critical level r_t^{min}

$$r_{t+1}^l < r_t^d - \frac{(1 + r_t^d) E_t}{A_t} \equiv r_t^{min}$$

The regulator requires the bank to maintain a constant default probability of $1 - \alpha$, so that $r_t^{min} \leq r_{t+1}^l$ with probability α . Assuming a normal i.i.d. probability distribution for r_{t+1}^l with mean $E(r_{t+1}^l)$ and standard deviation σ , we can write this value-at-risk (VaR) constraint as

$$E(r_t^{min}) = E(r_{t+1}^l) - \Phi\sigma \leq r_{t+1}^l \text{ with probability } \alpha$$

where Φ is the (constant) distance to default expressed as a multiple of the standard deviation of r^l (Shin (2010)).

The bank is assumed to be risk-neutral and thus maximizes expected profit. It is then easy to see that the banks' problem would be unbounded if it was not for the VaR constraint. Hence, the VaR constraint holds with equality

and the banks' optimal leverage is

$$\text{LVG}_t \equiv \frac{A_t}{E_t} = \frac{(1 + r_t^d)}{\Phi\sigma - E(r_{t+1}^l - r^d)}$$

which is equation (4) in the paper. This equation implies that banks will seek to implement a constant leverage over time – given the expected return and risk of their portfolio and given their refinancing rates.

As in Shin (2010) we assume that banks stabilize leverage through adjustments in debt and not via issuing or buying back equity. Adrian and Shin (2010b) show that this assumption is in line with the actual behavior of U.S. banks. Negative return shocks lead banks to reduce leverage by lowering debt given equity. Positive return shocks induce them to take on more debt.

Let us now think of σ as a bank (-type) specific parameter that reflects the bank's degree of geographical diversification. Consider two (types of) banks, one, called I , that is geographically diversified (integrated) and one local bank, L , that is working only in one state. Then, $\sigma^L > \sigma^I$ and the integrated bank will have higher leverage. Also, for a given decline in the deposit rate r^d , the leverage of the integrated bank increases by more than that of the local bank:

$$\frac{\partial \text{LVG}}{\partial r^d} = \frac{(\Phi\sigma - \mu) - (1 + r^d)}{(\Phi\sigma_n - \mu)^2} = \frac{\Phi\sigma_n - E(r_{t+1}^l) - 1}{(\Phi\sigma - \mu)^2} = \frac{1}{(\Phi\sigma_n - \mu)} - \frac{1}{(\Phi\sigma - \mu)^2} < 0$$

where $\mu = E(r_{t+1}^I - r^d)$ denotes the “term spread” or excess return on risky assets. This expression decreases (increases in absolute value) as σ_n decreases.¹⁹

Equilibrium in the U.S. money market and global imbalances. In what follows, we think of the U.S. banking system as being composed of banks of the I and the L type. Then equilibrium in the money market implies

$$D_t^{US} = \frac{E_t^I(1 + r^d)}{\Phi\sigma^I - E(r_{t+1}^I - r_t^d)} + \frac{E_t^L(1 + r^d)}{\Phi\sigma^L - E(r_{t+1}^L - r_t^d)} - (E_t^I + E_t^L)$$

where D_t^{US} is the supply of deposits that, for our purposes here, we take as exogenous. We can rewrite this condition as

$$\frac{D_t^{US}}{(E_t^I + E_t^L)} = \frac{(1 + r_t^d)\lambda_t}{\Phi\sigma^I - E(r_{t+1}^I - r_t^d)} + \frac{(1 + r_t^d)(1 - \lambda_t)}{\Phi\sigma^L - E(r_{t+1}^L - r_t^d)} - 1 = L_t^I\lambda_t + L_t^L(1 - \lambda_t) - 1$$

where

$$\lambda_t = \frac{E_t^I}{E_t^I + E_t^L}$$

can be interpreted as the degree of regional integration of the U.S. financial system (defined as the share of all equity owned by regionally integrated banks). Now consider an exogenous shock ΔD_t^{US} to deposit supply – the emergence of global imbalances. With constant total U.S. banking equity

19. Note that the term is negative as long as $\Phi\sigma_n - E(r_{t+1}^I) - 1 < 0$, which will always be the case, since otherwise, there would be a positive probability of the expected gross minimum return $1 + E(r^{min}) = 1 + E(r_{t+1}^I) - \Phi\sigma_n$ being negative, which is ruled out by the bank’s limited liability.

$E^{US} = E_t^I + E_t^L$ and constant regional diversification λ we get:

$$\frac{\Delta D_{t+1}^{US}}{E_t^I + E_t^L} = \left(\frac{\partial \text{LVG}^I}{\partial r^d} \lambda_t + \frac{\partial \text{LVG}^L}{\partial r^d} (1 - \lambda_t) \right) \Delta r_t^d$$

Here, we can think of $\Delta D_{t+1}^{US} / (E_t^I + E_t^L)$ as the theoretical counterpart of our capital inflow measures, CAPFLOW_t . Hence, the increase in capital inflows translates into an offsetting decline in interest rates. As we have seen, a given interest rate decline will affect leverage — and thus lending — more for I -banks than for L -banks. Hence, at the level of individual states, we will also see that lending increases more strongly in integrated states than in segmented states with only local banks.

Endogenous geographical diversification. Once we allow banks to react to the savings glut by also changing their geographical diversification, the market clearing condition becomes:

$$\begin{aligned} \frac{\Delta D_{t+1}^{US}}{E_t^I + E_t^L} = & \left(\frac{\partial \text{LVG}^I}{\partial r^d} \lambda_t + \frac{\partial \text{LVG}^L}{\partial r^d} (1 - \lambda_t) \right) \Delta r_t^d \\ & + \left[(\text{LVG}^I - \text{LVG}^L) + \left(\frac{\partial \text{LVG}^I}{\partial \sigma^I} \times \frac{\partial \sigma^I}{\partial \lambda} - \frac{\partial \text{LVG}^L}{\partial \sigma^L} \times \frac{\partial \sigma^L}{\partial \lambda} \right) \right] \Delta \lambda_t \end{aligned}$$

Here, the first part of the second term, $\text{LVG}^I - \text{LVG}^L$, captures how an increase in λ increases the ability of the U.S. financial system to absorb additional deposits as local banks become I -banks, *given* each bank type's respective level of geographical diversification (σ^I and σ^L respectively). The second part of the term captures, how an increase in λ may also change the geographical diversification of banks' portfolios. Clearly, it makes sense

to assume that regional diversification lowers notional portfolio risk, so that $\partial\sigma/\partial\lambda < 0$. Note also that

$$\frac{\partial \text{LVG}}{\partial \sigma_n} = \frac{-\Phi}{[\Phi\sigma_n - E(r_{t+1}^l - r_t^d)]^2} < 0$$

and that

$$\frac{\partial \text{LVG}}{\partial \sigma_n \partial r^d} = \frac{\Phi}{[\Phi\sigma_n - E(r_{t+1}^l - r^d)]^3} > 0$$

Hence, as refinancing rates r^d fall after the savings glut, the sensitivity of leverage to changes in the volatility of the portfolio increases. Thus, for any given marginal cost of lowering σ_n (which in the above comes about by increasing λ), banks will now have a stronger incentive to increase their geographical diversification, making diversification an endogenous reaction to the savings glut.

Appendix D: Appendix: Additional Results

Impact of capital inflows on aggregate U.S. house prices. We consider two scenarios: one in which the effect of capital inflows on the level of house prices is permanent and one in which it is transitory. Note that our estimate of the coefficient α in Table 2 captures the impact of capital inflows on house prices over the next four quarters. In both scenarios, we assume that the house price response builds up uniformly over four quarters to hit our estimate of the one-year ahead house price change as implied by α . In the permanent scenario, we assume that the house price response levels off after these four quarters. In the transitory scenario, we assume instead that the effect on house prices is uniformly undone over quarters 5 to 8. Hence, quarter-on-quarter house price growth in each state is a moving average of capital inflows of the form

$$\frac{HP_t^s}{HP_{t-1}^s} = \left(1 + \delta_1 \sum_{h=1}^4 \text{OPEN}^s \times \text{CAPFLOW}_{t-h} + \delta_2 \sum_{h=5}^8 \text{OPEN}^s \times \text{CAPFLOW}_{t-h} \right)^{1/4}$$

where HP_t^s is the index of house prices in state s normalized to 100 in the fourth quarter of 1996. The coefficients are given by $\delta_1 = (1 + \alpha)^{1/4} - 1$ in both scenarios and $\delta_2 = 0$ ($\delta_2 = -\delta_1$) in the permanent (transitory) scenario. We take the values of α from the two most demanding specifications in Table 2 reported in columns (4) and (5). We choose the higher of the two ($\alpha = 0.32$) for the permanent and the lower ($\alpha = 0.26$) for the transitory scenario. We use the above formula to generate predicted house prices for each state and then average across states. The results are provided in Figure D.1.

State-level indicators of geographical diversification of banks. As shown in Appendix B, aggregating individual banks' lending decisions to the state-level

implies that a de facto openness measure should have the general form

$$\text{OPEN}_{t-1}^s = \sum_{n=1}^N \lambda_{n,t} \omega_{n,t-1}^s$$

where the parameter $\lambda_{n,t} \in [0, 1]$ captures the extent to which bank n is diversified across states and $\omega_{n,t-1}^s$ is the share of bank n in total bank mortgage lending in state s at time $t - 1$ and N denotes the number of banks.

To construct $\lambda_{n,t}$ and $\omega_{n,t-1}^s$, we obtain data from the call reports published by the Federal Reserve Bank of Chicago over the period 1984 to 1995. For each bank, we then identify whether it is affiliated with a bank holding company that owns banks also in other states. If it is, we call it an integrated bank, otherwise we call it a local bank. We then construct four different versions of OPEN_{t-1}^s . The first is just a dummy indicating if a bank is integrated or not. In this case, OPEN_{t-1}^s equals the interstate asset ratio proposed by Morgan et al. (2004). Second, analogous to the construction of the interstate asset ratio, we use a dummy indicating if a bank belongs to a BHC that holds mortgage assets in another state. This gives rise to what we call the interstate mortgage ratio.²⁰ Third, we use the number of states in which the BHC to which an integrated bank belongs is active as an indicator of bank-level diversification. To obtain a measure of $\lambda_{n,t}$ between zero and one, we divide this number by the number of states in our sample. Our fourth measure of $\lambda_{n,t}$ is the Herfindahl index of a BHC's asset holdings across states.

20. In this case, for consistency, we use the share of a banks' mortgage lending in all state-level mortgages as the weights instead of total assets.

Alternative measures of capital inflows, net vs. gross inflows. Table D.2 presents results for alternative measures of capital inflows. Columns (1)–(2) of Table D.2 show that, while our results hold for gross inflows as well, they are stronger for net inflows. In fact, once we include both gross inflows and outflows (interacted with $OPEN^S$) in the same regression, the two coefficients are almost equal but of opposite sign, suggesting that net inflows are the relevant variable here. Our theory model allows us to discuss in detail why we believe net inflows to provide a sharper link with house prices.

The current account could misrepresent actual inflows into the U.S. money and capital markets for various reasons. First, it neglects valuation changes on foreign asset holdings. Clearly, such valuation changes could impact demand and supply for credit in the mortgage market by affecting private household wealth and the balance sheets of financial intermediaries. In column (3), following Favilukis et al. (2012), we therefore look at the change in the net holding of U.S. securities owned by foreigners as an alternative measure of capital inflows. Different from the current account, this variable takes on board potential valuation effects and also excludes foreign direct investment inflows which we would not expect to have a direct impact on the supply of dollar liquidity and the mortgage market.

Arguably, a large share of U.S. capital inflows over the late 1990s and early 2000s was motivated by global demand for U.S. safe assets. In column (4), we therefore focus on the change in foreigners' net holdings of *safe* U.S. securities, defined here as changes in the holdings of U.S. government bonds and mortgage-backed securities issued by government sponsored enterprises such as Freddie Mac and Fannie Mae. In both specifications our results hold.

The banking glut. Borio and Disyatat (2011) and Shin (2012) have argued that to understand the vulnerabilities that had built up in the financial sector in the years before the financial crisis, it is important to consider gross investment positions of foreign banks in the U.S.. Before 2008, U.S. affiliates (subsidiaries and branches) of foreign banks borrowed heavily in the U.S. money market. At the same time, the foreign parents of these affiliates built considerable long-term positions in the U.S. mortgage market. Shin (2012) calls this feature of global imbalances the banking glut (as opposed to the savings glut) and argues that it played a major role in the excessive risk taking in the U.S. financial sector by effectively enhancing the intermediation capacity of the U.S. financial system. When short-term dollar financing dried up in 2008, the balance sheets of international banks operating in the U.S. were therefore extremely vulnerable and became a major factor in the international transmission of the crisis. This transmission was so forceful because it got amplified through high leverage (large gross positions) even though the net position of foreign banks vis-à-vis the U.S. was actually quite small.

The extension of our model to international banks presented in the main text allows us to discuss how net capital inflows may be related to the banking glut. As documented by Shin (2012), falling refinancing rates in the U.S. interbanking markets encouraged European banks to borrow in U.S. dollars and to invest the funds into the U.S. housing market. However, European banks did not generally originate mortgage loans themselves. They usually bought securities backed by a geographically diversified portfolio of mortgages from U.S. financial intermediaries. In so doing, foreign banks effectively mimicked the behavior of regionally diversified U.S. banks. In the

model, this would amount to a shift from E^F to E^I for given total banking equity, E , and thus an increase in λ and the effective share of equity of all U.S. based banks that is used to underpin the risks associated with domestic lending $(E^I + E^L)/E$. Hence, the banking glut increases the capital base of the domestic part of the U.S. financial system $(E^I + E^L)$ against which leverage can be taken. This should dampen the effects of net inflows on banks' refinancing rates, leverage, lending and house prices that we document in this paper.

In Table D.3, we examine the possibility that the build-up in international banking sector positions contributed to house price increases and that they did so more strongly in states that were financially more open. We focus in five measures of the banking glut: the sum of all claims of foreign banks on U.S. assets as well as the banks' net and the gross positions vis-à-vis the U.S. as well as changes of total claims and net holdings. All measures are normalized with U.S. GDP.

Of all specifications, only the change in the net position of foreign banks is strongly significant and positive individually, which would be consistent with the argument of Borio and Disyatat (2011) and Shin (2012). The other measures, total claims and gross positions are actually negatively significant if considered alone. This would be consistent with our model, in which the banking glut would actually increase the de facto risk taking capacity of the U.S. financial system. However, all banking glut measures are insignificant once we also include our baseline measure of capital inflows, $CAPFLOW_t$. By contrast, the size of the coefficient on our baseline measure and its significance remain unchanged relative to our earlier specifications.

TABLE D.1. House Prices and de facto financial openness

Dependent variable is house price growth , $HPGROWTH_t^s$				
	$OPEN^s = open_{defacto}^s$			
	interstate asset ratio (1)	interstate mortgage ratio (2)	$\frac{1}{HHI}$ (3)	number of states (4)
$OPEN^s \times CAPFLOW_t$	0.530 (2.45)	0.628 (2.94)	0.240 (2.09)	0.090 (2.48)
CONTROLS	YES	YES	YES	YES
Observations	2820	2820	2820	2820
R^2	0.73	0.74	0.73	0.74

The Table shows the results from the panel regression

$$HPGROWTH_t^s = \alpha \times OPEN^s \times CAPFLOW_t + \text{CONTROLS} + \tau_t + \delta^s + \epsilon_t^s,$$

where $OPEN^s$ is defined in column headers and $CAPFLOW$ are capital inflows, measured as the negative current account balance over lagged total U.S. banking equity ($CAPFLOW_t = -CA_t / BankEquity_{t-1}$). The dependent variable, $HPGROWTH_t^s$, is the four-quarter ahead change in house prices. **CONTROLS** comprises the full set of controls from Table 2: **BASELINE**, $CREDIT \times OPEN^s$, and $PRE97^s \times CAPFLOW_t$. The vector **BASELINE** includes the growth rates of per capita state personal income, and population growth, the house price income ratio and the inter-state-branching indicator by Rice and Strahan (2010). The vector $CREDIT_t$ includes variables capturing aggregate credit demand and supply: the real treasury constant maturity rate, the real 3-months T-bill rate, the Baa-Aaa (default) spread, the survey of lending attitudes of senior loan officers and the Michigan survey of consumer house buying conditions index. The vector $PRE97^s$ includes 1985-1996 means of the following state-level variables: house price growth, the growth rates of tradable and non-tradable GDP and wages, volatility of GDP growth, employment growth, the unemployment rate, real estate lending growth, exports relative to GDP and a measure of import exposure (Autor et al. (2013)) aggregated to the state-level.

Data are quarterly, sample period is 1997-2012. Sample includes 47 US states (Alaska, Hawaii, Delaware as well as District of Columbia are left out). OLS estimates, all regressions include time- and state-fixed effects. Standard errors are clustered by state and time, t-statistics appear in parentheses.

TABLE D.2. Alternative measures of capital inflows, net vs. gross inflows

Dependent variable is house price growth , $HPGROWTH_t^s$				
	(1)	(2)	(3)	(4)
	INFLOW _t		Net foreign holdings of US assets	Net foreign holdings of US securities
$OPEN^k \times CAPFLOW_t$	0.005 (2.26)	0.009 (1.84)	0.037 (2.17)	0.131 (2.93)
$OPEN^s \times OUTFLOW_t$		-0.006 (-1.17)		
CONTROLS	YES	YES	YES	YES
Observations	2820	2820	2820	2820
R^2	0.70	0.70	0.70	0.70

The Table shows the results from the panel regression

$$HPGROWTH_t^s = \alpha \times OPEN^s \times CAPFLOW_t + \text{CONTROLS} + \tau_t + \delta_s + \epsilon_{s,t},$$

where $OPEN^s = 1995 - \text{Year of Banking Deregulation}$. The dependent variable, $HPGROWTH_t^s$, is the four-quarter ahead growth in house prices. $CAPFLOW_t$ is represented by alternative measures of capital inflows, which vary from column (1) to (4). In column (1) $CAPFLOW_t$ is measured by financial inflows over lagged total U.S. banking equity. In column (2), additionally to financial inflows financial outflows over lagged total U.S. banking equity are considered. In column (3) $CAPFLOW_t$ is defined as net foreign holdings of total assets over GDP, in column (4) as net foreign holdings of total securities over GDP. **CONTROLS** comprises the full set of controls from Table 2: **BASELINE**, $CREDIT \times OPEN^s$, and $PRE97^s \times CAPFLOW_t$. **BASELINE** includes the growth rates of per capita state personal income, and population growth, the house price income ratio and the inter-state-branching indicator by Rice and Strahan (2010). The vector $CREDIT_t$ includes variables capturing aggregate credit demand and supply: the real treasury constant maturity rate, the real 3-months T-bill rate, the Baa-Aaa (default) spread, the survey of lending attitudes of senior loan officers and the Michigan survey of consumer house buying conditions index. The vector $PRE97^s$ includes 1985-1996 means of the following state-level variables: house price growth, the growth rates of tradable and non-tradable GDP and wages, volatility of GDP growth, employment growth, the unemployment rate, real estate lending growth, exports relative to GDP and a measure of import exposure (Autor et al. (2013)) aggregated to the state-level. Data are quarterly, sample period is 1997-2012, sample includes 47 US states (Alaska, Hawaii, Delaware and District of Columbia are left out). All regressions are estimated by OLS and include time- and state-fixed effects. T-statistics are in parentheses, standard errors are clustered by state and time.

TABLE D.3. Banking glut vs. CAPFLOW

	Dependent variable is HPGROWTH _t ²					
	(1)	(2)	(3)	(4)	(5)	(6)
	total banking claims	net banking holdings	gross banking holdings	Measure of <i>BANKFLOW</i> is changes in claims	changes in net banking holdings	changes in gross banking holdings
$OPEN^S \times CAPFLOW_t$	0.029 (2.58)	0.026 (1.81)	0.028 (2.53)	0.031 (2.53)	0.032 (2.74)	0.030 (2.54)
$OPEN^S \times BANKFLOW_t$	-0.007 (-1.99)	0.020 (2.38)	-0.005 (-2.29)	0.013 (0.94)	-0.006 (-0.31)	0.009 (1.60)
CONTROLS	yes	yes	yes	yes	yes	yes
R^2	0.72	0.71	0.72	0.70	0.72	0.70

The Table shows the results from the panel regression

$$HPGROWTH_t^2 = \alpha \times OPEN^S \times CAPFLOW_t + \alpha_{BC} \times OPEN^S \times BANKFLOW_t + \text{CONTROLS} + \tau_t + \delta_s + \epsilon_{s,t},$$

where $OPEN^S = 1995 - \text{Year of Banking Deregulation}$ and $CAPFLOW$ are capital inflows, measured as the negative current account balance over lagged total U.S. banking equity ($CAPFLOW_t = -CA_t / BankEquity_{t-1}$). The dependent variable, $HPGROWTH_t^2$, is the four-quarter ahead change in house prices. The various measures of banking flows, $BANKFLOW_t$, are defined in the column headings. Sample period is 1997:2012, quarterly data. The sample includes 47 US States (Alaska, Hawaii, Delaware and District of Columbia are left out). **CONTROLS** are defined as in Table 2 column 4. All regressions include time- and state-fixed effects, t-statistics in parentheses are based on standard errors that are two-way clustered.

TABLE D.4. Placebo-sample: institution-level regressions for non-banks

	dep. variable: β_t^b			dep. variable: ∇_t^b		
	σ_n at level			σ_n at level		
	county	MSA	State	county	MSA	State
	(1)	(2)	(3)	(4)	(5)	(6)
$RTC_{t-1} \times CAPFLOW_t$	0.002 (0.128)	0.112 (0.317)	0.360 (0.776)	0.000 (0.025)	0.008 (0.099)	0.037 (0.105)
log total assets	0.039 (0.835)	0.041 (0.832)	0.061 (0.895)			
Bank Asset Growth	0.000 (0.409)	0.000 (0.365)	0.000 (0.236)			
RTC_{t-1}	-0.018 (-0.929)	-0.100 (-0.266)	-0.753 (-1.396)	-0.000 (-0.025)	-0.007 (-0.096)	-0.035 (-0.055)
Observations	9,204	9,204	9,204	9,204	9,204	9,204
R ²	0.024	0.025	0.025	0.000	0.000	0.000

The first three columns of the table present institution-level regressions of the lending supply shocks β_t on bank characteristics and aggregate factors of the form

$$\beta_t^b = \alpha \times RTC_{t-1}^b \times CAPFLOW_t + \text{CONTROLS} + \delta^b + \tau_t + \nabla_t^b$$

on the placebo-sample of non-bank financial institutions. In the regression **CONTROLS** includes the logarithm of bank total assets, aggregate lending growth and the interaction $RTC_{t-1} \times \text{CREDIT}_t$ with **CREDIT** the vector of aggregate credit supply factors of Favilukis et al. (2012). Bank's risk taking capacity RTC_{t-1} is computed according to equation (10), with σ_{t-1} calculated according to equation (12) at the county (column 1), MSA (column 2) and state levels (column 3) respectively, $E(r^l - r^d) = r^d = 0.01$ and $\Phi = 2.32$.

The right panel (columns 3-6) presents the regressions of the residual ∇_t^b of the respective regression in column 1-3 (re-run with $RTC_{t-1} \times CAPFLOW_t$ dropped) on $RTC_{t-1}^b \times CAPFLOW_t$, RTC_{t-1} and bank and time fixed effects. All regressions are WLS using lagged U.S. wide market share of banks as weight. Standard errors are clustered by bank and year. Numbers in parentheses are t-statistics. Sample period is 1997-2005.

TABLE D.5. Bank-level regressions with alternative capital flow measures

	dep. variable: β_t^b		dep. variable: ∇_t^b	
	(1)	(2)	(3)	(4)
$RTC_{t-1} \times INFLOW_t$	4.852 (2.927)	8.057 (3.628)	0.180 (1.176)	0.065 (0.151)
$RTC_{t-1} \times OUTFLOW_t$		-4.761 (-4.291)		-0.064 (-0.121)
log total assets	0.015 (0.030)	0.030 (0.103)		
Bank Asset Growth	0.003 (0.636)	0.003 (0.868)		
RTC_{t-1}	2.345 (2.175)	0.588 (0.900)	-0.266 (-0.472)	-0.054 (-0.337)
Observations	31,732	31,732	31,732	31,732
R ²	0.081	0.080	0.000	0.000

The first two columns of the table present bank-level regressions of the bank lending supply shocks β_t on bank characteristics and aggregate factors of the form

$$\beta_t^b = \alpha_1 \times RTC_{t-1}^b \times INFLOW_t + \alpha_2 \times RTC_{t-1}^b \times OUTFLOW_t + \text{CONTROLS} + \delta^b + \tau_t + \nabla_t^b$$

where **CONTROLS** includes the logarithm of bank total assets, aggregate lending growth and the interaction $RTC_{t-1} \times CREDIT_t$ with **CREDIT** the vector of aggregate credit supply factors of Favilukis et al. (2012). Bank's risk taking capacity RTC_{t-1} is computed according to equation (10), with σ_{t-1} calculated according to equation (12) at the MSA level, $E(r^l - r^d) = r^d = 0.01$ and $\Phi = 2.32$.

The right panel (columns 2-4) presents the regressions of the residual ∇_t^b of the respective regression in column 1-2 (re-run with capital in- and outflows dropped) on $RTC_{t-1}^b \times INFLOW_t$, $RTC_{t-1}^b \times OUTFLOW_t$, RTC_{t-1} and bank and time fixed effects. All regressions are WLS using the (lagged) U.S.wide market share of banks as weight. Standard errors are clustered by year of liberalization ($OPEN^s$) of the home state of the bank and by year. Numbers in parentheses are t-statistics. Sample includes all commercial banks, 1997-2005.

TABLE D.6. List of MSAs straddling state borders

	MSA Name	States
1	Allentown-Bethlehem-Easton	PA-NJ
2	Augusta-Richmond County	GA-SC
3	Boston-Cambridge-Newton	MA-NH
4	Cape Girardeau	MO-IL
5	Charlotte-Concord-Gastonia	NC-SC
6	Chattanooga	TN-GA
7	Chicago-Naperville-Elgin	IL-IN-WI
8	Cincinnati	OH-KY-IN
9	Clarksville	TN-KY
10	Columbus	GA-AL
11	Cumberland	MD-WV
12	Davenport-Moline-Rock Island	IA-IL
13	Duluth	MN-WI
14	Evansville	IN-KY
15	Fargo	ND-MN
16	Fayetteville-Springdale-Rogers	AR-MO
17	Fort Smith	AR-OK
18	Grand Forks	ND-MN
19	Hagerstown-Martinsburg	MD-WV
20	Huntington-Ashland	WV-KY-OH
21	Kansas City	MO-KS
22	Kingsport-Bristol-Bristol	TN-VA
23	La Crosse-Onalaska	WI-MN
24	Lewiston	ID-WA
25	Logan	UT-ID
26	Louisville/Jefferson County	KY-IN
27	Memphis	TN-MS-AR
28	Minneapolis-St. Paul-Bloomington	MN-WI
29	Myrtle Beach-Conway-North Myrtle Beach	SC-NC
30	New York-Newark-Jersey City	NY-NJ-PA
31	Omaha-Council Bluffs	NE-IA
32	Philadelphia-Camden-Wilmington	PA-NJ-DE-MD
33	Portland-Vancouver-Hillsboro	OR-WA
34	Providence-Warwick	RI-MA
35	Salisbury	MD-DE
36	Sioux City	IA-NE-SD
37	South Bend-Mishawaka	IN-MI
38	St. Joseph	MO-KS
39	St. Louis	MO-IL
40	Texarkana	TX-AR
41	Virginia Beach-Norfolk-Newport News	VA-NC
42	Washington-Arlington-Alexandria	DC-VA-MD-WV
43	Weirton-Steubenville	WV-OH
44	Wheeling	WV-OH
45	Worcester	MA-CT
46	Youngstown-Warren-Boardman	OH-PA

FIGURE D.1. Capital inflows and house prices —quantitative importance for the average U.S. State

NOTES: The solid line plots the cross-state average level of house prices in the United States (unweighted mean, 1996=100). The dotted line is the level of house price for the average state that is explained by capital inflows according to our estimates in Table 2, column (4), assuming that the impact of capital inflows on the level of house prices is permanent and builds up uniformly over four quarters. The dashed line gives the average level of house prices explained by the model according to the estimates in Table 2, column (5) and assuming that the impact of capital inflows on the level of house prices is transitory, i.e. first builds up over four quarters and then falls to zero over the next four quarters. For details see the main text of the paper.

FIGURE D.2. Years since liberalization and geographical diversification of banks

NOTES: The figures plot our baseline openness measure—the number of years passed since interstate banking deregulation—against various *de facto* measures of openness: the 1986–1993 average of the interstate asset ratio, the share of mortgages in a state issued by integrated banks, the inverse of the Herfindahl Index, $1/HFI^s$, and the average number of states in which banks in a state are operating. Cross-sectional regressions of four *de facto* measures on the baselines measure, $OPEN_{de\,facto}^s = \beta OPEN^s + constant + \epsilon_s$, yield the following coefficient estimates (t-statistics) and R^2 : interstate asset ratio: 0.04 (4.51) $R^2=0.30$; interstate mortgage ratio: 0.03 (4.21) $R^2=0.27$, inverse of Herfindahl: 0.05 (3.22) $R^2=0.17$; Average number of states in which states banks are operating: 0.13 (2.19) $R^2=0.08$.

NOTES: The figures plot our baseline openness measure —the number of years passed since interstate banking deregulation — against various banking characteristics measures: the 1986-1993 average of the interstate asset ratio, the share of mortgages in a state issued by integrated banks, the inverse of the Herfindahl Index, $1/HFI^s$, and the average number of states in which banks in a state are operating.

FIGURE D.4. A stylized model of state-level bank lending

NOTES: The figure illustrates how local banks (on the left) and integrated banks (on the right) react to a positive loan supply shock (savings glut/ refinancing shock) (upper panel) and a loan demand shock (lower panel). In both panels, the supply curve of integrated bank is flatter (and more curved) than that of the local bank because due to its geographical diversification it can take on more leverage for any given level (and increase) of the interest rate.

A loan supply shock — represented here by a drop in the refinancing cost Δr — shifts down banks' loan supply curve and tilts it to the right. Lending increases and lending rates decline and the effect is particularly strong for the integrated bank, leading it to lower its lending rates and to increase its lending more than the local bank. The term spread of lending over refinancing rates will decrease in particular for the integrated bank. Over time, some local bank customers will migrate to the integrated bank which charges lower rates until both banks charge the same mortgage rate again (the move from the solid to the dashed demand curve).

A positive loan demand shock (shift from D to D') increases lending rates of both integrated and local banks. Again, the increase in lending is larger for the integrated bank, but this time refinancing rates stay constant for both banks and term spreads rise. Since the initial increase in rates is higher for the local bank, some of its customers move away towards the integrated bank, equalizing rates between banks over time (demand curve D'').