FINANCIAL INTEGRATION, HOUSING AND ECONOMIC VOLATILITY

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Abstract

The Financial Crisis and the Great Recession illustrate the sensitivity of the economy to a housing bust. This paper shows that financial integration, fostered by deregulation allowing banks to form nationwide branch networks, amplified housing-price volatility and increased the economy's sensitivity to local housing-price shocks. We exploit variation in credit-supply subsidies across local markets from the Government-Sponsored Enterprises to measure housing price changes unrelated to fundamentals. Using this instrument, we find that a 1% rise in housing prices causes a 0.25% increase in economic growth. This effect is larger in localities more financially integrated with other markets through bank ownership ties. Financial integration thus raised the effect of collateral shocks on the economy, thereby increasing economic volatility.

I. INTRODUCTION

The recent 'Great Recession', many argue, had its origins in the boom and bust in housing, and the knock-on effects of the resulting financial crisis (Brunnermeier, 2008). Some argue that the length and depth of this recession stems from the slow recovery of housing and the associated debt overhang for consumers (Mian and Sufi, 2011). In this paper, we study links from housing to the overall economy in the years leading up to the crash (1994 to 2006). During this period, local housing prices became more volatile as regions such as the Sun Belt experienced dramatic booms. Figure 1 plots the mean absolute growth shock of local housing prices from 1975 to 2006. Volatility trends down during the 1970s and 1980s. Starting in the 1990s, however, volatility stops falling and then begins to rise. This trend break coincides with changes in the financial and banking systems in the US, which have become increasingly well integrated as deregulation allowed banks to form nationwide branch networks and as securitization allowed mortgage credit to flow easily across markets. We show that shocks to local housing demand were amplified by financial integration because capital could flow freely across connected markets. Financial integration also strengthened the link from housing to the overall economy.

Financial integration may dampen or amplify economic shocks. Morgan, Rime and Strahan (2004) – MRS hereafter – show theoretically that integration's effect on volatility depends on the sources and magnitudes of shocks hitting the local economy. With integration, local economies become more insulated from shocks to the supply of local finance (e.g. local bank capital). During the 1980s and early 1990s, these shocks were a major source of business-cycle instability (Bernanke and Lown, 1991). For example, the number of bank and S&L failures during the 1980s averages more than 150 per year (Kroszner and Strahan, 2008), and the collapse of the S&L industry amplified downturns in areas such as Texas and California. Integration makes local economies less sensitive to these financial disturbances because capital can flow in from external sources and thus allow investment to continue, even if local lenders are distressed. MRS show empirically that state-level banking integration fostered by deregulation during the 1970s and 80s lowered volatility of local economies in these years.

MRS's theoretical model, however, also shows that integration, by allowing financial capital to flow away from depressed areas and into booming ones, can amplify local cycles. For example, if collateral values rise sharply in a locality, borrower debt capacity and demand for credit increases; integration helps bring financial resources from abroad to satisfy higher credit demand. The influx of credit from external sources raises growth above what would have been possible in a stand-alone, or dis-integrated, financial system. These flows correspondingly reduce collateral values from areas with relatively weak credit demand because these market face capital outflows. Thus, capital flows generated by credit demand shocks will reduce co-movements in collateral values across financially integrated markets.

Beyond its effects on capital flows, integration is also associated with lower investment by lenders in private information about local business conditions, borrower credit quality and housing-price fundamentals (Loutskina and Strahan, 2011; Romero-Cortes, 2011). As a result of securitization, for example, residential mortgage credit supply responds more now to changes in the market value of collateral than in the past because lenders condition their credit decisions more on public signals (e.g. borrower FICO scores and loan-to-value ratios) and less on private information (Rajan, Seru and Vig, 2010). Both of these forces – more 'flighty' capital and more reliance on public information – may increase collateral volatility and raise the sensitivity of local cycles to variation in collateral values. Consistent with these ideas, we find that financial integration during our sample raises the volatility of housing prices, that shocks in the housing sector have a quantitatively substantial causal impact on local economies, and that the transmission of these housing-price shocks increases with financial integration.

The analysis proceeds in three steps. First, we document a positive relationship between financial integration and the magnitude of local house-price shocks. To do so, we measure financial integration at the level of the Central Business Statistical Area (CBSA), the US Census Bureau's definition of a city. The measure (*In-CBSA ratio)* is based on the ownership of bank branches across CBSAs, equal to the fraction of local deposits owned by a banking company also owning branches in other CBSA markets. So, a CBSA in which all of its branches are owned by banks with branches in other CBSAs would have *In-CBSA ratio* = 100%.

We find that the volatility of shocks to CBSA-level housing price growth increases with financial integration. The effect increases in magnitude when we use variation across states in restrictions on interstate branching as an instrument for financial integration (Rice and Strahan, 2010). Thus, there is a robust difference in local house-price volatility between more- and less-integrated local markets. This result reverses that of MRS, who use data from the 1970s and 1980s, when shocks to the financial sector were an important source of business-cycle variation.[1] Our results, however, are consistent with the theoretical argument that, in the absence of shocks to financial institutions, integration amplifies the impact of collateral shocks. To test this mechanism, we compare shocks for all unique pairs of local markets. If integration increases capital flightiness in response to collateral values shocks, then integration between pairs of markets ought to reduce the correlation between shocks across markets. Using housing price

[1] Like MRS, we have also tested whether the amount of deposits in external markets, as a second integration measure, affects volatility. This second integration measure is also positively related to volatility in some specifications, although its magnitude is smaller and less significant than our primary integration measure.

changes to proxy for such demand shocks, we find this result. Markets that are *more* integrated with each other have *less* similar changes in housing prices, controlling for trends (time dummies), for pair-wise fixed effects and for the similarity of industry composition. Again, we find that the effects increase in magnitude when we instrument for integration using a pair-wise combination of each area's regulatory stance toward interstate branching.[2]

In the second part of the analysis, we build an instrument for house-price appreciation that exploits the importance of the Government-Sponsored Enterprises (GSEs) – Fannie Mae and Freddie Mac – in housing finance. Fannie and Freddie subsidize mortgage credit, but only for mortgages that fall below the jumbo-loan threshold (Loutskina and Strahan, 2009). Borrowers with housing demand near the jumbo-loan threshold stand to benefit from an increase in the threshold, leading to an increase in housing demand and housing prices (Adelino, Schoar and Severino, 2011). While the jumbo-loan cutoff changes uniformly across CBSAs, its effects vary across markets. For example, in Los Angeles - where about 5.3% of mortgages were made to borrowers within 5% of the jumbo-loan cutoff - the change in cut-off would have a bigger impact than in Wichita, Kansas - where this fraction was about 0.5%. Since there is both cross-sectional and time-series variation in the amount of such demand (e.g. LA v. Wichita), we generate a set of instruments based on the product of the sensitivity to changes in the jumbo-loan cutoff in market i during year $t-1$ times the change in the cutoff itself between years $t-1$ and t. The instruments depend only on the distribution of mortgage credit during the preceding year and the change in the jumbo-loan cutoff during the current year, which is the same across all local markets and depends mechanically on lags of increases in nationwide prices. Furthermore, we exploit the elasticity of the housing supply across different geographies to better capture the

[2] Kalemni, Papaionnou and Peydro (2010) find similar effects following financial integration across 20 developed economies.

response of housing prices to changes in demand (Saiz, 2010). Thus, it is plausible to assume that these instruments pick up variation in changes in housing demand exogenous to overall economic fundamental in the local area.

We find that these instruments are powerful. Local housing prices appreciate faster in markets where credit on jumbo borrowers was more constrained in the prior year, based on the distribution of borrowers around the jumbo cutoff. This effect is stronger in markets with relatively inelastic housing supply because prices are more sensitive to changes in demand where the physical supply of housing is limited by geographic barriers.

Armed with exogenous variation in housing prices, the third part of the analysis shows that housing prices have a strong causal impact on local economic growth in employment and output. In our base model, a 1% increase in housing prices causes an increase in local GDP growth of about 0.25% and an increase in non-construction, non-finance employment growth of about 0.15%. The latter effect implies that higher prices spill over to sectors not directly affected by housing. We then show that the effects of house-price shocks are stronger in local markets with high levels of financial integration than in markets with low integration. In local areas one-standard deviation above the mean level of financial integration, a 1% housing price shock leads to a 0.30% increase in GDP growth. Taken together – higher housing price volatility and increased sensitivity to house-price shocks – the results imply that financial integration has increased economic volatility, both by amplifying variation in collateral values (house prices) and by strengthening links from collateral to the overall economy.

Our paper contributes to three strands of the literature. First, the effect of financial integration on economic volatility has been explored both across US states and also in the context of liberalization of international capital markets (e.g., Morgan, Rime and Strahan, 2004; Demyanyk, Ostergaard and Sorenson, 2007; Kalemni, Papaionnou and Peydro (2010)). We find that integration can amplify shocks and de-sychronize asset markets in an environment of strong credit demand and a profitable financial sector. In other settings, where financial shocks are important, integration can increase synchronization because credit supply shocks propagate across connected markets (e.g. Peek and Rosengren, 2000). Second, conventional explanations for the US housing boom blame loose lending practices as a key driver of price appreciation (e.g., Mian and Sufi (2009), Keys et al (2010), Demyanyk and Van Hemert (2010), Loutskina and Strahan (2011)). Yet these studies do little to explain why booms were concentrated in places like as Florida, Arizona and California. Financial integration can rationalize regional booms by allowing capital to flow into areas with strong credit demand.

Third, many have argued that the so-called 'Great Recession' has its root in the crash of housing prices beginning in the middle of 2006. Our results are consistent with this explanation but also suggest that the economic boom was itself fueled by house-price appreciation. The findings extend the work of Mian and Sufi (2009 and 2011), who show that household debt and consumption were strongly correlated with house-price appreciation during the boom. Conversely, declines in consumer spending and financial distress across local markets during the bust are also associated with declines in housing equity. Unlike Mian and Sufi (2011), however, we go a step further and estimate the total effect of housing price shocks on the economy, and we condition this estimate on aspects of the financial system. Shocks to housing have had a large effect on the overall economy, especially in markets that are well integrated nationally.

In the next section we briefly review the forces leading to increased integration over time. In Section III, we describe our integration measures in detail, and document their link to local volatility. Section IV then estimates the relationship between shocks to housing prices and local growth. Here, we first establish a first-stage model that relates changes in credit-supply subsidies from the GSEs to house-price appreciation. We then use this model to generate an instrument for housing price changes to estimate its causal impact on the economy as a whole. Section V concludes.

II. FORCES OF CHANGE LEADING TO FINANCIAL INTEGRATION

Deregulation integrates the banking system

Into the 1970s, most lending occurred through insured depository institutions, and technological, legal and regulatory barriers prevented integration across geographical and product markets. Over time, these barriers have eroded. The process began during the 1970s, when only 12 states allowed unrestricted statewide branching and another 16 prohibited branching entirely. Between 1970 and 1994, 38 states eased their restrictions on in-state branching. States also prohibited ownership of their banks by out-of-state bank holding companies. These barriers to integration began to fall when Maine passed a 1978 law allowing entry by out-of-state BHCs if, in return, banks from Maine were allowed to enter those states. Other states followed suit, and state deregulation of intra-state banking was nearly complete by 1992 (Jayaratne and Strahan, 1996; Jayaratne and Strahan, 1998).

The transition to full *inter*state banking and branching was fostered by passage of the Interstate Banking and Branching Efficiency Act of 1994 (IBBEA), which effectively permitted bank holding companies to enter other states without permission and allowed banks to operate branches across state lines (Rice and Strahan, 2010). With these legal changes, banks now operate across many states and localities, which allows financial resources to flow more easily across geographical markets through banks' internal capital markets (Houston, James and Marcus, 1997).

Despite the passage of IBBEA, states continue to exercise authority under this law to restrict or limit interstate branch entry. While IBBEA opened the door to nationwide branching, it allowed states to influence the manner in which it was implemented. States that opposed entry by out-of-state banks could use provisions of IBBEA to erect barriers to some forms of out-of-state entry, to raise the cost of entry, and to distort the means of entry. From the time of enactment in 1994 until the branching default "trigger date" of June 1, 1997, IBBEA allowed states to employ various means to erect these barriers. States could set regulations on interstate branching with regard to four important provisions: (1) the minimum age of the target institution, (2) whether or not to permit *de novo* interstate branching, (3) whether or not to permit acquisition of individual branches rather than whole banks, and (4) how tightly to control the percentage of deposits in insured depository institutions controlled by any single bank or bank holding company. Following Rice and Strahan (2010), we use these four state powers to build a simple index of interstate branching restrictions across states. The index equals zero for states that are most open to out-of-state entry. We add one to the index when a state adds any of the four barriers just described. Specifically, we add one to the index: if a state imposes a minimum age on target institutions of interstate acquirers of 3 or more years; if a state does not permit *de novo* interstate branching; if a state does not permit the acquisition of individual branches by an out-of-state bank; and if a state imposes a deposit cap less that 30%. So, the index ranges from zero to four. We use this index below as a policy instrument to help explain variation in our key measure of financial integration.

Securitization integrates housing finance

The move toward integration in mortgage lending occurred in concert with branching deregulation, initially spurred by the activities of the Government-Sponsored Enterprises (GSEs) - The Federal National Mortgage Association (Fannie Mae) and the Federal Home Loan Mortgage Corporation (Freddie Mac). By the 1990s, both Fannie Mae and Freddie Mac had become heavy buyers of mortgages from all types of lenders, with the aim of holding some of those loans and securitizing the rest. Together they have played the dominant role in fostering the development of the mortgage secondary market. As shown by Frame and White (2005), the GSEs combined market share has grown rapidly since the early 1980s. In 1990 about 25% of the $2.9 trillion in outstanding mortgages were either purchased and held or purchased and securitized by the two major GSEs. By 2003, this market share had increased to 47%.[3] This market share fell after 2004 in the wake of the accounting scandals and the growth of subprime mortgages by private lenders, and then increased significantly since 2006 in response to the credit crisis. GSE access to implicit government support allows them to borrow at rates below those available to private banks, and to offer credit guarantees on better terms than competitors without such implicit support.[4]

As shown in Loutskina and Strahan (2010), the GSEs enhance mortgage liquidity, reduce the cost of borrowing, and increase mortgage acceptance rates conditional on borrower credit

[3] GNMA provides a very important source of mortgage finance to low-income borrowers, holding or securitizing about 10% of all mortgages outstanding.

[4] Passmore, Sherlund and Burgess (2005) argue that most (but not all) of the benefits of GSE subsidies accrue to their shareholders rather than mortgage borrowers. To take advantage their low borrowing costs, during the 1990s the GSEs increasingly opted to hold, rather than securitize, many of the mortgages that they buy. Policymakers became concerned about the resulting expansion of interest rate risk at the GSEs (Greenspan, 2004), although the 2008 crisis resulted more from the credit guarantees offered by the agencies than from exposure to their retained mortgage portfolio.

quality. The GSEs buy and hold some mortgages, and they also often securitize them. When the GSEs buy mortgages, they bear both credit and interest rate risk. When GSEs securitize mortgages, they either buy them and issue mortgage-backed securities (MBS), or they just sell credit protection to the original lender. In the first case, the originating bank retains no stake in the mortgage. In the second case, the bank continues to fund the mortgage and bear the interest rate risk, but obtains the option to sell the mortgage off as an MBS (because of the credit protection). In all cases, the GSEs enhance liquidity and thus foster integration of credit markets.

The GSEs operate under a special charter, however, that limits the size of mortgages that they may purchase or securitize. These limitations were designed to ensure that the GSEs meet the legislative goal of promoting access to mortgage credit for low and moderate-income households. The GSEs may only purchase non-jumbo mortgages, defined in 2006 as those below $417,000 for loans secured by single-family homes. The loan limit increases each year by the percentage change in the national average of single-family housing prices during the prior year, based on a survey of major lenders by the Federal Housing Finance Board. The limit is 50% higher in Alaska and Hawaii. Because the loan limit changes mechanically and only as a function of *national* housing prices, local housing supply or demand conditions have no effect on the jumbo loan cutoff. We exploit this fact in developing our instrument for housing price growth below.

Starting in the early 1980s, securitization moved beyond the GSEs, as private investment banks began to purchase and securitize jumbo loans, providing similar services for large mortgages that Fannie and Freddie provide for non-jumbos, although without the government subsidy. This fact can be seen by the jump in the average mortgage interest rate around the jumbo loan cutoff. In fact, this rate differential increased sharply during the financial crisis as lenders became more reluctant to extend credit and more constrained in their ability to finance their investments.

Both the moves to allow geographical expansion of banks within and across states, as well as the expansion of GSEs and private securitization have benefited both lenders and borrowers. Diversification opportunities have been enhanced, credit can flow more easily toward high-return investments, and opening up of markets has increased competition and lowered the price of credit (Demsetz and Strahan, 1997; Stiroh and Strahan, 2003). As we show next, however, financial integration has led to greater volatility, both in the housing sector (and, by extension, for the value of collateral more generally) and also in the economy as a whole.

III. FINANCIAL INTEGRATION AND HOUSE-PRICE VOLATILITY

In this section we test how financial integration affects the volatility of housing prices within local markets, and how the synchronicity (or interrelatedness) of housing price changes between markets varies with pair-wise measures of financial integration. In our first set of models, we build a panel dataset based on house-price volatility and financial integration at the level of the Central Business Statistical Area (CBSA) over the 1994 to 2006 period (unit of analysis = CBSA-year). In the second set of models, we build a richer panel by creating all CBSA-year pairs, again over the 1994 to 2006 period (unit of analysis = CBSA-pair-year). We test whether the correlation or similarity of housing prices shocks between pairs of markets changes as the two markets become more financially integrated with each other.

To start, we measure the volatility of the housing prices using the absolute deviation of housing price growth in a CBSA-year from the conditional mean, after removing time and CBSA fixed effects. Specifically, we estimate the following regression:

$$Ln\ Housing\ Price_{i,t} - Ln\ Housing\ Price_{i,t-1} = \alpha_t + \gamma_i + growth\text{-}shock_{i,t}\ . \qquad (1)$$

Data for housing price growth rates are constructed from the Federal Housing Finance Association's (FHFA) CBSA-level house price index. The residual $growth\text{-}shock_{i,t}$ captures how much housing prices growth differs in each CBSA and year compared to average housing price growth in this year across all geographies. The absolute value of this residual reflects housing price fluctuations specific to a given geography: $Vol_{i,t}=|growth\text{-}shock_{i,t}|$.

The CBSA-year regressions test how integration affects housing-price volatility, as follows:

$$Vol_{i,t} = \alpha_t + \gamma_i + \beta^1 Integration_{i,t} + Other\ Controls + \varepsilon_{i,t}\ , \qquad (2)$$

where $Integration_{i,t}$ equals our measures of the extent to which financial activity in a CBSA-year is connected to financial activity in other CBSAs (defined below).

The pair-wise regressions have the following structure:

$$Interrelatedness_{i,j,t} = \alpha_t + \gamma_{i,j} + \beta^2 Integration_{i,j,t} + Other\ Controls + \varepsilon_{i,j,t} \qquad (3a)$$

where $Interrelatedness_{i,j,t}$ equals the negative of the absolute value of the difference in housing-price growth shocks between two CBSAs in a given year:

$$Interrelatedness_{i,j,t} = -\left| growth\text{-}shock_{i,t} - growth\text{-}shock_{j,t} \right| \qquad (3b)$$

So, an increase in $Interrelatedness_{i,j,t}$ measures a decline in the difference in growth shocks between two CBSAs. In Equation (3a), $Integration_{i,j,t}$ measures the pair-wise connectedness of two CBSA markets in a given year (defined below).

As noted in the introduction, financial integration may raise volatility either because integrated lenders condition their credit decisions more on prices and less on other dimensions of credit risk (e.g. specialized knowledge about the local economy), or because capital flows more easily toward high-demand markets and away from low-demand markets. Both channels imply $\beta^1 > 0$ in Equation (2). By looking at integration's effects on pair-wise markets, we can isolate the capital flows channel. Imagine two CBSA markets – 'A' and 'B' – that are well integrated. A shock to prices in 'A' (and thus to credit demand there) will draw financial resources away from 'B', thus accommodating the credit demand and raising prices in A and lowering them in B. This second capital flight channel thus suggests that financial integration ought to make house-price changes become less correlated as integration between two markets increases, so $\beta^2 < 0$ in equation (3a).[5]

Measuring Financial Integration by CBSA-year

Our measure of financial integration is built from the distribution and ownership of bank branches and deposits across local markets. The measure is based on information on total deposits, location and ownership of all bank branches from the Federal Deposit Insurance Corporation's (FDIC) *Summary of Deposits*, available online annually from 1994 forward.[6] We construct the *In-CBSA ratio,* equal to the fraction of all deposits in a CBSA that are owned by a holding company which also owns deposits in one or more other CBSAs.[7]

[5] House price variation driven by local credit supply shocks will tend to attenuate this effect.

[6] See http://www2.fdic.gov/sod/.

[7] We define a banking company as the highest entity within a bank holding company for banks owned by holding companies, or for the bank itself for stand-alone banks.

Variation in the *In-CBSA ratio*$_{i,t}$ depends on bank entry decisions into market *i* in year *t*, which in turn may reflect risk management or diversification motivations of potential entrants. Since the intrinsic volatility of a particular market may play a role in this entry decision, the relationships observed in the fixed effects OLS estimate of Equation (2) could be biased by reverse causality. For example, if out-of-state banks prefer to enter safe markets, the coefficient on financial integration would tend to be biased downward in OLS. To eliminate this potential source of bias, we also estimate Eq. (2) using an instrumental variable model, where the instrument for the *In-CBSA* ratio equals the index of restrictions on interstate branching described in Section II. This index ranges from zero to four, where four represents the highest level of barriers to entry by out-of-state banks. Since this index varies mainly across states, rather than within states over time, we do not have strong identification in the fixed effects model. Hence, we report OLS with and without CBSA fixed effects but only report the IV model without these effects. (All models include time effects.)

Measuring Integration by CBSA-year pairs

To measure integration between pairs of CBSAs, we build the *Common CBSA Ratio*. For each CBSA pair, we sum up all deposits with a common ownership link, add these across the two markets, and then divide by the total amount of deposits in the two CBSAs. Higher values of *Common CBSA Ratio* indicate a greater degree of shared financial resources – greater integration – between CBSAs. We also estimate our model with a dummy-variable version of *Common CBSA Ratio,* equal to one when there are any commonly owned deposits and zero otherwise. This second approach is arguably more robust than the first, and its coefficient is also somewhat easier to interpret. As already mentioned, since bank entry decisions may be endogenously driven by economic conditions in local markets, we use an instrument for the

Common CBSA Ratio, again based on the state-level branching restrictions index. In this case, since each observation represents a pair of CBSAs, the instrument equals the sum of the branching restrictions index in the states where the two CBSAs are located. Hence, we again report both the fixed effects OLS model as well as the IV model.[8]

Table 1 reports summary statistics for our volatility and integration measures. Panel A reports the CBSA-year level means and standard deviations for house-price volatility and the four integration measures; Panel (B) reports these statistics for the two pair-wise interrelatedness measures and the pair-wise integration measure. The *In-CBSA ratio* average 81.4% (Panel A), indicating that in the typical CBSA-year the majority of deposits are owned by banking companies with deposits elsewhere. This variable has substantial variation – mainly in the cross section – with a standard deviation of 15.3%. The average house-price growth shock equals 4.56%, suggesting substantial CBSA-specific shocks to local markets after removing trends in overall housing price appreciation. The pair-wise data tell a similar story, with an average difference in growth residuals between pairs of CBSAs of 4.07 percentage points. Almost 40% of market pairs have some ownership links, with an average *Common CBSA ratio* of 8.28%.

Volatility increases with integration

Table 2 reports our estimation of Equation (2), linking financial integration to total house-price growth volatility, along with the first-stage model for the *In-CBSA Ratio*. All models include time fixed effects to take out aggregate trends as well as the national business cycle. In addition, we control in all models for the share of employment across the following different industry segments: construction, mining and logging; finance; education and health

[8] We include the pair-wise fixed effects even in the IV model. Since the instrument depends on branching in two areas rather than one, a change in the branching index in *either* locality's state generates within-CBSA variation over time. Thus, we get strong identification in the first-stage model, even including the pair-wise fixed effects.

services; manufacturing; trade, transportation and utilities; information technology; professional and business services and other services.[9] In some models, we also incorporate CBSA-level fixed effects to capture time invariant market-level characteristics that may be correlated with volatility. In every case, we cluster data at the CBSA level to build our standard errors.

The results strongly suggest, first, that financial integration is greater in CBSAs located in states with fewer restrictions on interstate branching (Table 2, column 1). An increase in the branching index from 0 to 4 – from least to most restrictive – comes with a decline in the *In-CBSA ratio* of about 5%, which is large relative to the variation in this variable (σ = 15.3% - see Table 1). The branching restrictions index has strong explanatory power in the first stage as well, with a t-statistic above 3. The F-statistic for the first stage regression equation is 10.3, which confirms that we have a well-identified model and pass the weak instruments test of Stock and Yogo (2005).

Second, financial integration is associated with greater volatility of housing prices (columns 2 & 3). *In-CBSA ratio* has a positive and significant effect on volatility in OLS without the CBSA effects (column 2); in this OLS model, however, the economic magnitude of integration is small. As noted, however, endogenous entry by banks may bias the coefficient on integration downward (that is, toward zero), and this notion is supported by the IV model, where the coefficient rises in magnitude substantially. In this model (column 3), a standard deviation

[9] The industry share variables are built off the industry employment numbers provided by the Bureau of Labor Statistics. The employment data is provided at detailed industry level. We aggregate the data at the level of 9 different industries: (i) construction, mining and logging; (ii) manufacturing including durable and non-durable goods manufacturing; (iii) trade, transportation, and utilities; (iv) information; (v) financial activities; (vi) professional and business services; (vii) education and health services; (viii) leisure and hospitality; and (ix) Other services. For each industry, we computer the percentage contribution to the CBSA level employment. The employment in the government sector is the omitted variables.

increase in the *In-CBSA ratio* would increase house-price growth volatility by 0.4%, a substantial increase relative to the dispersion in house-price volatility (σ = 2.8% - see Table 1).

Interrelatedness across markets falls with integration

Table 3 reports the estimation of Equation (3a), along with the first stage model linking integration between pairs of CBSA markets (*Common CBSA ratio*) to the sum of the branching restrictions index in the two states. In these pair-wise models, the dependent variable equals the negative of the absolute value of the difference in house-price growth shocks in a given year (recall Equation (3b) above). As noted, all of the models include time fixed effects and a separate fixed effect for every unique pair of CBSAs – a total of 65,508 unique fixed effects. These fixed effects remove factors such as geographical distance that may affect the similarity of housing markets between two CBSAs. We also include a variable capturing the 'distance' or similarity of the industry mix between pairs, equal to the sum of squared difference in industry shares (i.e. the Euclidean distance). This pair-wise factor will capture variation over time in the differences in industry mix between markets. We also group our data into clusters for each CBSA to build standard errors. So, although the models are built from nearly one million observations, there are just 362 independent clusters.

Table 3 reports the results for specifications using the continuous measure of integration (*Common CBSA ratio* = the fraction of commonly owned deposits), and using a dummy variable equal to one for markets that have some degree of commonly owned bank deposits. The latter model is somewhat easier to interpret and also may be more robust to outliers. Columns (1) and (2) report the first stage models, where for both the continuous and dummy variable approaches we have very strong identification (t-stat > 10 for the branching restrictions instrument). For example, increasing the degree to which a CBSA pair are restricted from cross ownership from 0 (most open) to 8 (least open) would come with a 16% increase in the probability that the two CBSAs have some common ownership in deposits (column 2).

Consistent with Table 2, we find that markets that are more integrated with each other have less commonality in growth shocks, and we also find that magnitudes increase when we instrument for integration with branching restrictiveness (columns 3-6). For example, the indicator variable model suggests that markets that share bank deposits have house-price growth shocks that are 4.4% less similar, which is large relative to the overall variation of these differential shocks ($\sigma = 4.13\%$ - see Table 1). The results support the idea that capital flows affect collateral values. In markets that are financially connected, markets with high credit demand (e.g. high house prices) can draw on financial capital from markets with lower demand, thereby reducing the correlatedness of collateral values between the two markets. In markets that share financial resources, housing price growth rates become *less similar*. This result is strong evidence that financial integration amplifies credit-demand shocks; capital flowing between these markets lowers the similarity in shocks to the value of collateral.

IV. THE EFFECTS OF HOUSING PRICES ON ECONOMIC GROWTH

In this section we ask two questions. First, did the increase in housing-price volatility lead to greater business-cycle instability? Second, did financial integration strengthen the link from housing prices to overall economic performance, thus further raising overall volatility? The first question is motivated by the trend toward greater housing price volatility (recall Figure 1). The second question is suggested by theories of financial integration, which imply that more mobile financial capital should strengthen the link from shocks to credit demand – e.g. housing prices, or more generally, the value of collateral – and economic output.

To answer these questions, we trace out the causal impact of shocks to housing prices on overall economic output by CBSA-year ($Y_{i,t}$), measured by personal income growth, employment growth, employment growth without sectors directly affected by housing (construction and finance) and GDP growth. Specifically, we estimate panel regressions with the following structure:

$$Y_{i,t} = \alpha^{y}_{t} + \gamma^{y}_{i} + \beta^{y}_{1}\, \textit{House-Price Growth}_{i,t} + \textit{Other Control Variables} + \varepsilon_{i,t} \tag{4a}$$

and

$$\begin{aligned} Y_{i,t} = {} & \alpha^{y}_{t} + \gamma^{y}_{i} + \beta^{y}_{1}\, \textit{House-Price Growth}_{i,t} + \beta^{y}_{2}\, \textit{Financial Integration}_{i,t} \\ & + \beta^{y}_{3} \textit{Financial Integration}_{i,t} * \textit{House-Price Growth}_{i,t} + \textit{Other Control Variables} + \varepsilon_{i,t}\,. \end{aligned} \tag{4b}$$

We estimate Equations (4a) and (4b) for our CBSA-year panel dataset from 1994 to 2006, including both year and CBSA fixed effects. The year effects remove trends as well as the national business cycles, while the CBSA effects take out long-run differences in average economic growth rates.

To test how financial integration affects links from house price shocks (or, more generally, collateral shocks), we interact *House-Price Growth*$_{i,t}$ with *In-CBSA ratio*, using the branching restrictions index as the instrument for *In-CBSA ratio*, as in Table 2. If changes in housing prices raise borrower debt capacity and, in turn, raises consumer demand and firm investment, then $\beta^{y}_{1} > 0$ (4a); if financial integration, by allowing capital to flow in from external markets, strengthens this effect, then $\beta^{y}_{3} > 0$ in (4b). In order to estimate the overall impact of housing on the economy, we first estimate Equation (4a) without financial integration, and then estimate models with the interaction term in (4b).

As additional controls variables, we include the share of employment across industry sectors as before; three measures of the strength and health of the local banking sector: the average capital-asset ratio, the log asset size of banks operating in the CBSA, and the average growth rate of assets of local banks; and, in some specifications, one lag of the dependent variable.[10]

GSE Housing-Finance Subsidies as a Source of Instruments for Housing Price Growth

Shocks to the overall economy will both affect and be affected by the value of housing, as well as the value of real estate and collateral more generally. Our aim is to trace out the causal impact of shocks to housing on the overall economy; hence, we need instruments that move housing prices (and so are sufficiently powerful) but otherwise remain unrelated to fundamental drivers of economic growth (and so meet the exclusion restriction for valid instruments). We use subsidies in housing-finance from the GSEs to build such instruments.[11] Potential home buyers receive a financing subsidy through the activities of the GSEs, who stand ready to buy mortgages that fall below the jumbo-loan cutoff and meet a set of credit-worthiness underwriting criteria.[12] The cut-off is binding on borrowers, as is evident from the histogram of loan applications and loan approval rates presented in Figures 2A and 2B (adapted from Loutskina and Strahan (2009)). The large spike in loan applications and approval rates just

[10] Industry shares are from the Bureau of Labor Statistics. Bank characteristics are taken from the Bank *Call Reports*; CBSA-level averages equal the weighted average of banks operating in the CBSA based on the share of deposits held in a given CBSA by each bank.

[11] Adelino, Schoar and Severino (2011) use a similar strategy at the transaction level to trace out how GSE subsidies affect the price per square foot of housing.

[12] For evidence about the size of this subsidy, see Loutskina and Strahan (2010).

below the jumbo cut-off indicate that the funding is both more abundant and cheaper below the jumbo loan cut-off. The cutoff is the same everywhere (except Alaska and Hawaii), and it increases annually based on a mechanical formula linked to changes in national housing prices. The increase in the jumbo-loan cutoff thus raises the subsidy to some potential home buyers, but the increase, crucially, is not dependent on conditions in the local area (CBSA).

We exploit the idea that the impact of this increased subsidy varies across local housing markets. For example, in a market where all home prices fall below the jumbo-loan cutoff in *t-1*, home buyers there would receive no incremental benefit from an increase in the cutoff in year *t*; all potential homebuyers would already be subsidized. In contrast, in markets with substantial demand near the jumbo-loan threshold, potential homebuyers would benefit greatly when the cutoff rises.

We use two strategies to measure differences across markets in the impact of changes in the jumbo-loan cutoff on housing demand. Detailed data for all mortgage applications to lenders above $50 million in assets are collected annually under the Home Mortgage Disclosure Act (HMDA). The HMDA data include loan size, whether or not a loan was accepted, some information on borrower credit characteristics, and the location of the property down to the Zip code level. Using these data, we estimate the fraction of loan applications in CBSA *i* and year *t-1* that are above the jumbo cutoff then, but would fall below that cutoff in the subsequent year (year *t)* as a consequence of the increase in the cutoff between the two years. This ratio captures the percentage of borrowers that would benefit from the change in the cut-off though getting access to more readily available and/or cheaper credit.

This first instrument is incomplete because it ignores the borrower self-selection into the area just below the cut-off (recall Figure 2A). A large fraction of home buyers reduce their borrowing to fall below the cut-off in year *t-1*, but many would also benefit from an increase in the jumbo-loan cutoff . For example, often home buyers will increase their equity investment in a property to be able to finance their borrowed funds in the subsidized, non-jumbo segment. Others will split their borrowing into a senior, non-jumbo mortgage (to gain the subsidy), and finance the remainder with a second-lien mortgage from a portfolio lender (i.e. a lender who holds the mortgage) plus equity. Thus, many mortgage applicants below – but not too far below – the jumbo-loan cutoff would also benefit from its increase. To capture this portion of demand, we build an instrument equal to the total fraction of applications within 5% of the jumbo-loan cutoff (on either side) in year *t-1*, multiplied by the percentage change in the cutoff between years *t-1* and *t*.

For each instrument, we also add an interaction with a measure of housing-supply elasticity built for 263 CBSAs based on physical impediments to expansion in the housing stock, such as waterways, mountains, and so on.[13] Saiz (2010) shows that cities with high supply elasticity have both slower increases in housing prices over time and faster population growth, compared to low-elasticity cities. These results make sense because low barriers to the expansion of housing implies that increased demand from population growth can be accommodated without increasing the cost of housing (e.g. land is not scarce in these areas). In our setting, we expect prices to respond more to the demand shocks associated with changes in the jumbo-loan cutoff in markets with low housing-supply elasticity than in markets with high elasticity.

[13] We use the elasticity estimates available online at: http://real.wharton.upenn.edu/~saiz/ and then convert them to the new definitions of CBSA using the zip-code overlap.

Figure 3 illustrates our identification strategy graphically for two extreme cases: a local market where most of the demand for housing is already subsidized by the GSEs and with very high supply elasticity (e.g. Wichita, where supply elasticity equals 5.5 and only 0.5% of total mortgage applications lie within 5 percentage points of the jumbo-loan cutoff), versus a market with a large mass of demand near the jumbo-loan cutoff and with low supply elasticity (e.g. Los Angeles, where supply elasticity equals 0.63 and about 5.4% of total mortgage applications lie within 5 percentage points of the jumbo-loan cutoff). An increase in the GSE jumbo-loan cutoff shifts housing demand only slightly in Wichita but substantially in Los Angeles. Because supply responds elastically in Wichita, prices barely rise. In LA, however, prices rise sharply, both because demand shifts further from the increased subsidy and because supply responds very little. Thus we trace a shock in a supply of funding to the housing price changes accounting for both CBSA-specific demand shifts and the CBSA-specific supply conditions.

The first-stage model then takes the following form:

$$\begin{aligned} \textit{House-Price Growth}_{i,t} = {} & \alpha^{HP}{}_{t} + \gamma^{HP}{}_{i} + + \textit{Other control variables} + \\ & + \beta_1{}^{HP}\textit{Share-New-NJ}_{i,t-1} + \beta_2{}^{HP}\textit{Share-New-NJ}_{i,t-1} \text{ x } \textit{Saiz-Elasticity}_i \\ & + \beta_3{}^{HP}\textit{Share-Near-NJ}_{i,t-1} + \beta_4{}^{HP}\textit{Share-Near-NJ}_{i,t-1} \text{ x } \textit{Saiz-Elasticity}_i + \varepsilon_{i,t}\,, \end{aligned} \tag{5}$$

where *Share-New-NJ*$_{i,t-1}$ equals the fraction of applications in CBSA *i* and year *t-1* that will fall below the jumbo-loan cutoff next year (year *t)*; *Share-Near-NJ*$_{i,t-1}$ equals the share of applications within +/- 5% of the cutoff in year *t-1* times the percentage change in the cutoff between *t* and *t-1*. We expect housing prices to grow fastest in markets with a large mass of demand that would benefit from an increase in the jumbo cutoff; thus, we expect: $\beta_1{}^{HP} > 0$, *and* $\beta_3{}^{HP} > 0$. Since house prices should react less if supply is elastic, we expect the interaction terms to offset, meaning $\beta_2{}^{HP} < 0$, $\beta_4{}^{HP} < 0$. We estimate Equation (5) with year and CBSA fixed effects, and we cluster the standard errors at the level of the CBSA. (Note that the direct effect of the Saiz elasticity measure, which is constant over time, is absorbed by the CBSA fixed effects.)

Results

Table 4 reports summary statistics for our instruments, for housing price growth and for personal income, employment and GDP growth during the 1994-2006 period. We obtain the CBSA-year level data on employment (and employment by segment) from the Bureau of Labor Statistics; the personal income data from the Bureau of Economic Analysis; and the local geography GDP from Moody's Analytics.[14]

The analysis begins in 1994 because the financial integration data, based on deposits, become available starting in 1994, and because HMDA data become available only in 1992. We end the analysis in 2006 for two reasons. First, we do not want our estimates to be driven by the Financial Crisis and the ensuing Great Recession. Second, our identification strategy relies on the consistent and mechanical increase in the jumbo-loan cutoff. This cutoff was raised aggressively and in response to political pressure during the Financial Crisis, and has subsequently remained fixed despite falling housing prices in an effort to bolster prices and sustain mortgage credit. The instrumental variables thus lose power after 2006, as they only generate an expansion in housing demand when the cutoff increases.

Table 5 reports the first-stage equation (Eq. (5)) linking the instruments to house-price appreciation, along with the time and CBSA fixed effects, industry share and banking sector control variables. We report the models first for each instrument separately (columns 1 and 3),

[14] The CBSA-year level GDP estimates are also available from Bureau of Economic Analysis (BEA) but only starting in 2001. We cross-reference the Moody's Analytics data with BEA and find the correlation of 98.7% between two data series.

we then report each instrument with its interaction with the Saiz supply-elasticity variable (columns 2 and 4), and last for all instruments together (column 5, which is the first stage regression used subsequently). All of the sets of instruments are powerful – with significant effects individually and collectively – although *Share-Near-NJ* is clearly stronger than *Share-New-NJ* (compare columns 1 and 3). Moreover, the signs and magnitudes of the coefficients are economically sensible individually. For example, a standard deviation increase in *Share-Near-NJ* leads to an increase in housing price growth of 2.7% (a little more than one-half of a standard deviation – see Table 4). Each instrument is also more positive in markets with low supply elasticity (columns 2 and 4). Sign patterns are difficult to interpret in the final regression, with both instruments and interaction terms, because the instruments are highly correlated (ρ=0.92). The F-test on all four instruments is 45.29 and passes the test for weak instruments with flying colors (Stock and Yogo, 2005).

Table 6 reports a baseline set of IV estimates linking the exogenous component of housing price appreciation to economic outcomes (Equation 4a). We estimate all models with time and CBSA fixed effects and with time-varying industry share variables, and time-varying measures of banking system characteristics. Table 6 reports a total of eight specifications - with and without the lagged dependent variable, times four different measures of output: personal income growth (columns 1 & 2), total employment growth (columns 3 & 4), the growth of total employment excluding employment in financial firms and construction (columns 5 & 6), and GDP growth (columns 7 & 8). Employment without construction and finance allows us to test whether any effects that we observe spillover beyond segments not directly tied to housing finance.

The coefficient estimates are statistically and economically significant across all specifications, ranging from 0.14 to 0.26. An exogenous 1% increase in housing prices (stemming from a credit supply increase) thus causes the local economy to expand by 0.14 to 0.26 percentage points faster than otherwise. The coefficients on total employment growth are smaller than GDP growth, which makes sense because GDP includes all sources of production from local sources (i.e. it includes returns to capital as well as labor).[15] Moreover, the coefficient on employment growth without segments directly tied to housing suggests that spillovers from higher collateral values raise output beyond the housing sector. Coefficients on personal income growth tend to be somewhat smaller because some of the variation depends on sources of income not tied specifically to the local area.[16]

Table 7 reports our last test, where we introduce an interaction between housing price growth and financial integration (Equation 4b). For this model, we add the branching restrictions index and its interaction with the other instruments and model housing price growth, financial integration and their interaction as endogenous variables.[17] Identification for the direct effect of financial integration is weak due to the inclusion of the CBSA fixed effects, but we are able to get strong identification for the interaction between housing prices and integration (since the interaction has both cross sectional and time series variation).

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[15] We have also estimated these models separately for the early (1994-2000) and late (2001-2006) portions of our sample. We find that housing is positively and statistically significantly related to economic outcomes in both samples, with somewhat larger magnitudes in the first half of the sample.

[16] We have explored several alternative ways to build instruments to check for model robustness. In one set of models, we only use the interaction between the Saiz elasticity measure with the share near non-jumbo * change in cutoff; these results are close to those reported in Table 6, both statistically and quantitatively. We have also estimated models in which we eliminated the time-variation in the share near non-jumbo by using its average value at the beginning of our sample. These results lead to somewhat larger coefficients on the house-price growth variable that have a higher level of statistical significance than those reported in Table 6.

[17] The branching index will also help identify housing growth, as Favara and Imbs (2010) show that housing prices grew faster in states more open to interstate banking due to greater availability of credit.

The results suggest that house price shocks have a greater impact on economic outcomes in financially integrated markets. Across all four specifications, housing price growth and financial integration are jointly significant at better than 1%. Moreover, the interaction term suggests that better integration has an economically important effect on the size of the causal impact of housing prices on economic output. For example, at the mean of the *In-CBSA ratio* (0.81), a 1% increase in housing prices would generate an increase in GDP growth of 0.15% (0.15 = -0.70+0.81*1.044); in markets one-standard deviation *above* the mean level of integration (0.81+0.15), the same 1% housing-price shock would lead to an increase of 0.30% (0.30 = -0.70+0.96*1.044). The interaction effect of integration on housing is statistically significant across all four models, with a magnitude that varies from 1.0 to 1.4. Because credit supply can respond more elastically to increases in collateral values when local markets are better integrated, an increase in housing prices generates a larger positive spillover in integrated markets. In these areas, the higher demand for credit can draw financial resources in from other sectors.

V. CONCLUSION

The Financial Crisis and subsequent Great Recession of 2007-2011 have emphasized for everyone the importance of a strong housing market to the economy. Housing markets not only increased sharply during the 2000s, but they also became more volatile across local markets. We show that this volatility increase is explained in part by better financial integration. We then demonstrate a causal link from housing to the overall economy, using variation in the impact of credit-supply subsidies from the GSEs to construct an instrument for housing price changes that is unrelated to economic conditions in the local economy. Our estimates suggest that a 1% rise in housing prices increase growth by about 0.25%. This effect is larger in localities that are better integrated with other markets through bank ownership ties. The results suggest that financial integration raises the effect of collateral shocks on the economy, thereby increasing economic volatility.

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Figure 1: Volatility of the Housing Prices

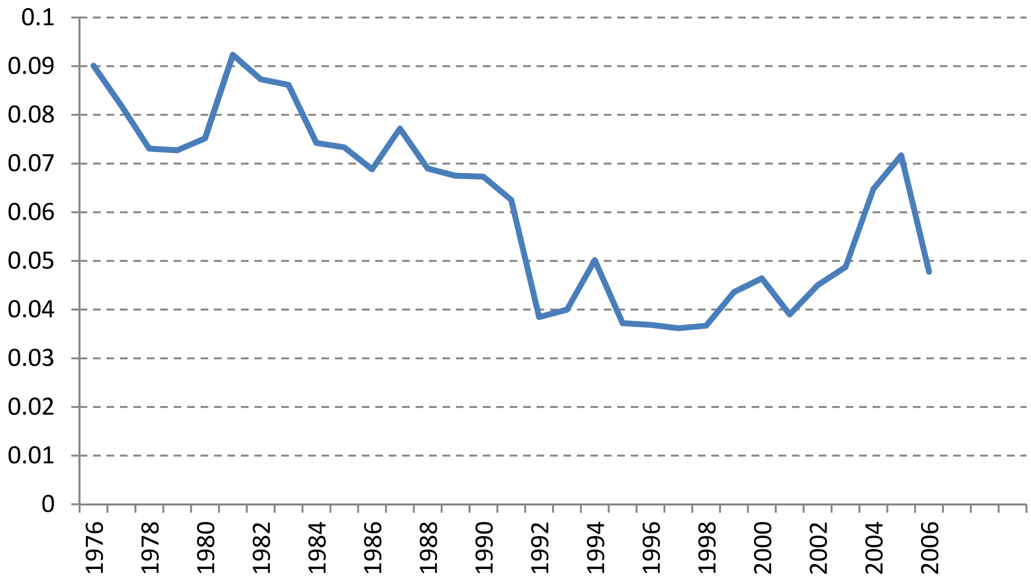


Figure 2A: Histogram of Loan Applications 1994-2006.

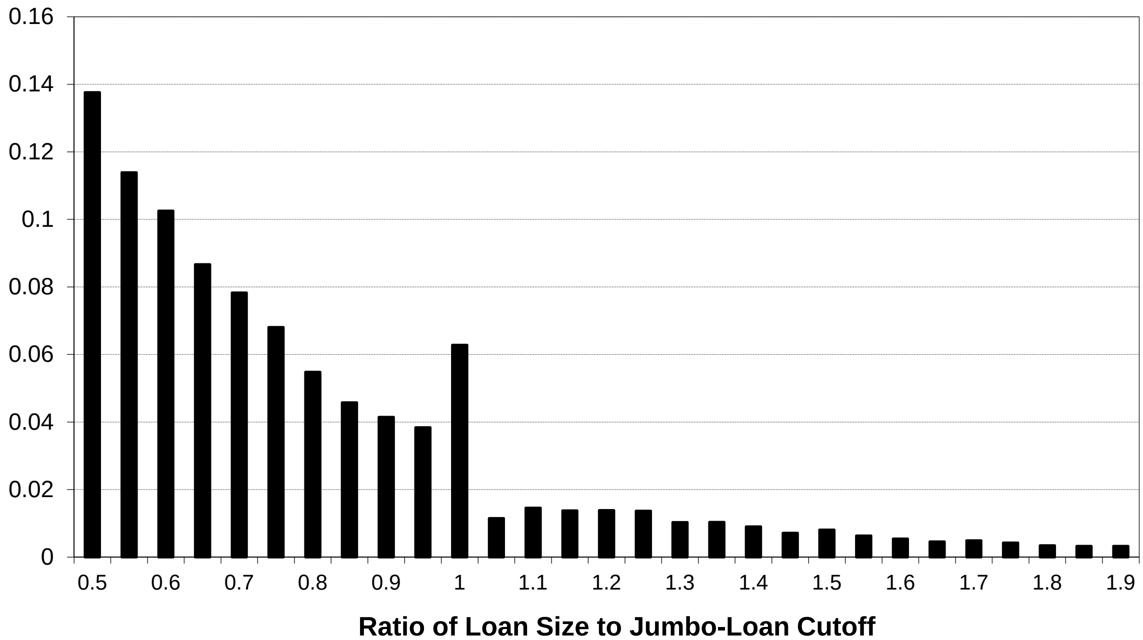


Figure 2B: Share of Approved Loan Applications 1994-2006.

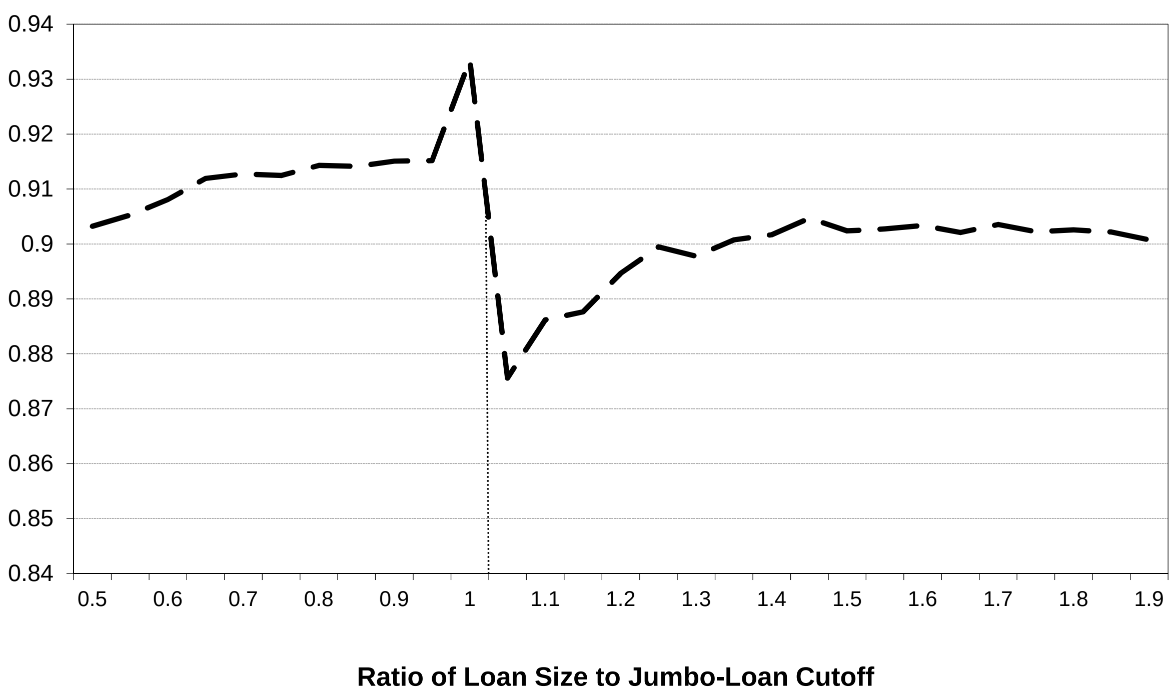
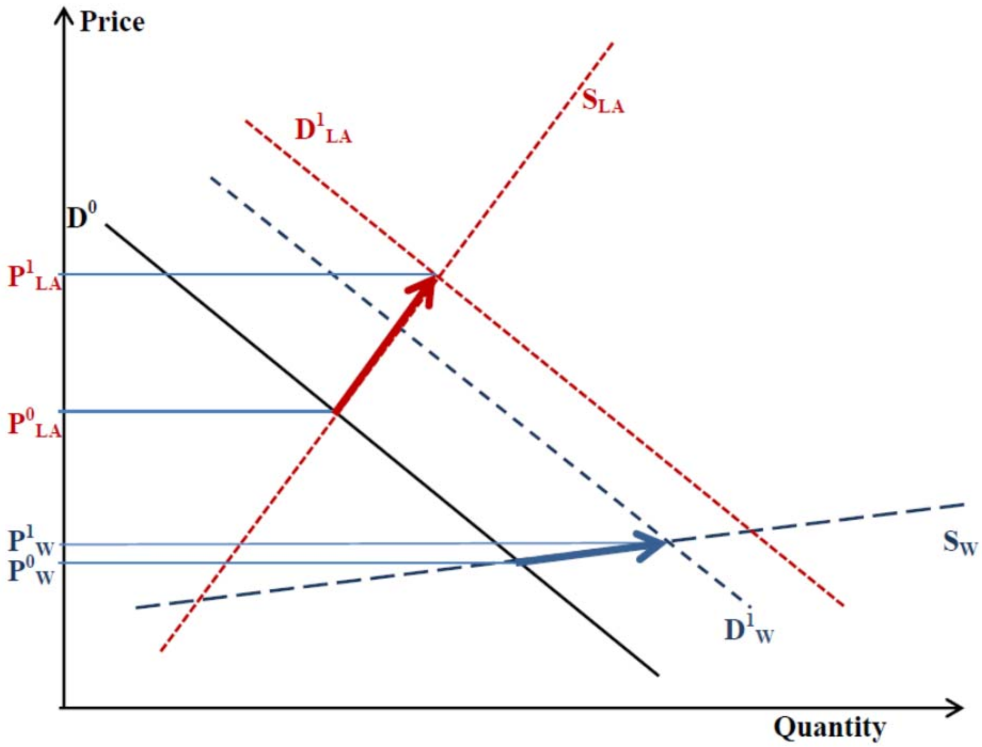


Figure 3: Responses of Different Markets to Changes in GSE Loan Cut-off



The graph illustrates the responses of two hypothetical markets to changes in the GSE loan cut-off. The subscript LA represents Los-Angeles CA and subscript W represents Wichita KS. Two markets are characterized by different elasticity of housing supply (S_{LA} and S_W) as well as different shifts in the demand curves caused by the same change in the loan cut-off (D^1_{LA} and D^1_W). The graph illustrates the corresponding changes in the housing prices.

Table 1: Summary Statistics for Measures of Integration and Housing Price Growth

Panel A reports summary statistics for two measures of financial integration that vary across CBSA-years. The In-CBSA ratio equals the fraction of deposits in CBSA-year that are owned by banking companies with deposits in other CBSAs. Panel B reports summary statistics at the level of CBSA-pair-years, where the measure of integration equals the sum of deposits with common ownership in a pair of CBSAs divided by total deposits in the two CBSAs.

Panel A: CBSA-Year Panel	***Mean***	***StDev***
In-CBSA Ratio	15.3%	81.4%
Housing Price Growth	4.55%	5.05%
Absolute Value of Housing Price Growth Residual	2.77%	4.56%
Panel B: CBSA-Pair-Year Panel		
% of shared deposits	8.28%	14.38%
% of shared deposits when positive	22.32%	16.03%
Indicator for CBSA pair with positve shared deposits	36.38%	N/A
- Absolute Value of Differential Growth Shock	-4.07%	4.13%

Table 2: Housing Price Volatility and Financial Integration

This table reports regressions of housing price volatility on measures of financial integration. The dependent variable is constructed as follows: first, we regress housing price growth on a CBSA fixed effect and year fixed effect and save the residual. We use the absolute value of this growth residual as teh dependent variable. Each model includes time effects. We report the OLS models with and without CBSA level fixed effects. The IV model is only well identified without the CBSA fixed effects. Standard errors are clustered by CBSA.

Dependent Variable:	In-CBSA Ratio	Absolute Value of Residual House-Price Growth		
	First-Stage	OLS	IV	OLS
	(1)	(2)	(3)	(4)
Branch Restriction Index	-0.0133***	-	-	-
	(3.02)	-	-	-
In-CBSA Ratio	-	0.00832**	0.0307**	0.00554
	-	(2.48)	(2.18)	(0.63)
Share of employment in construction, mining and logging	0.503**	-0.0199	-0.0298	-0.275***
	(2.27)	(-0.859)	(-1.164)	(-3.857)
Share of employment in financial sector	-0.898**	0.0735**	0.0941**	0.575***
	(2.22)	(2.30)	(2.34)	(3.46)
Share of employment in education and health services	-0.181	0.0319***	0.0351***	0.169*
	(1.45)	(3.70)	(3.68)	(1.94)
Share of employment in manufacturing	0.0544	0.0135**	0.0123*	-0.128*
	(0.52)	(2.03)	(1.67)	(-1.747)
Share of employment in trade, transportation, and utilities	0.0721	-0.00244	-0.00254	-0.0116
	(0.43)	(-0.255)	(-0.239)	(-0.170)
Share of employment in information	-0.164	-0.0098	0.00295	-0.303
	(0.21)	(-0.176)	(0.05)	(-1.216)
Share of employment in professional and business services	0.624***	-0.0236	-0.0387	-0.0725
	(3.27)	(-1.349)	(-1.539)	(-0.821)
Share of employment in leisure and hospitality	0.425***	-0.0143	-0.0226	0.272**
	(2.85)	(-1.157)	(-1.395)	(2.42)
Share of employment in other services	0.272	-0.0613	-0.0731	-0.372
	(0.50)	(-1.437)	(-1.520)	(-1.540)
Sum of Squared employment shares	-0.0381	0.00802	0.00713	-0.0302
	(0.13)	(0.47)	(0.37)	(-0.213)
Time Effects	yes	yes	yes	yes
CBSA Effects	no	no	no	yes
Number of Observations	4,397	4,397	4,397	4,397
R^2	10.0%	14.6%	13.4%	26.9%

*** p<0.01, ** p<0.05, * p<0.1

Table 3: Housing Price Interrelatedness Between Market Pairs and Financial Integration

This table reports regressions of the negative of the absolute value of the difference in housing price shocks between pairs of CBSA markets on measures of financial integration between the two market pairs. The dependent variable is constructed as follows: first, we regress housing price growth on a CBSA fixed effect and year fixed effect and save the residual. We use the absolute value of this growth residual as the growth shock in market i, year t. Each model includes time effects and CBSA-pair fixed effects. Standard errors are clustered by CBSA.

Dependent Variable:	Interrelatedness	Interelatedness Indicator	- Absolute Value of Differential Growth Shock			
	First-Stage	*First-Stage*	*OLS*	*OLS*	*IV*	*IV*
	(1)	(2)	(3)	(4)	(5)	(6)
Branch Restriction Index	-0.00432***	-0.0195***	-	-	-	-
	(10.41)	(10.65)	-	-	-	-
Interelatedness	-	-	-0.0245***	-	-0.200***	-
	-	-	(8.17)	-	(4.92)	-
Interelatedness Indicator	-	-	-	-0.00260***	-	-0.0442***
	-	-	-	(4.07)	-	(4.61)
Distance between Employment Shares	-0.00635	-0.0295	-0.0144**	-0.0143**	-0.0147**	-0.0147**
	(0.54)	(0.57)	(2.10)	(2.08)	(2.17)	(2.15)
Time Effects	yes	yes	yes	yes	yes	yes
CBSA-Pair Fixed Effects	yes	yes	yes	yes	yes	yes
Number of Observations	707,256	707,256	707,256	707,256	707,256	707,256
R^2	18.2%	20.2%	23.0%	23.0%	16.0%	14.0%

*** p<0.01, ** p<0.05, * p<0.1

Table 4: Summary Statistics for Economic Growth, Housing Price Growth and Instrument for Housing Price Growth

This table reports summary statistics for housing price growth, four measures of local economic growth, and two instruments built reflecting the distribtion of mortgage credit around the jumbo-mortgage cutoff.

	Mean	***StDev***
Housing Price Growth	5.41%	4.63%
Personal Income Growth	5.21%	2.55%
Employment Growth	1.46%	2.39%
Employment Growth, without construction and finance	1.14%	2.62%
CBSA level GDP growth	5.39%	3.04%
Share of New Non-Jumbo borrowers	0.357%	0.788%
Share Near the Jumbo Cutoff * Change in Cutoff	0.092%	0.145%
Saiz Measure of Housing Supply Elasticity	2.595	1.422

Table 5: Regressions relating Housing Price Growth to Distribution of Mortgage Credit around the Jumbo-Loan Cutoff

This table reports regressions of housing price growth by CBSA-Year on the share of borrowers in year t-1 that will become non-jumbo in year t (share new non-jumbo), and the total fraction of borrowers within +/- 5% of the jumbo-loan cutoff in year t-1 times the change in the jumbo loan cutoff between t-1 and t. All regressions include time and CBSA fixed effects, along with measures of industry structure and the health of the local banking system. Column 5 includes all instruments and acts at the first-stage for the subequent IV models (Tables 6 and 7). Standard errors are clustered by CBSA.

Dependent Variable:	*Housing Price Growth*					
	(1)	(2)	(3)	(4)	(5)	(6)
Share of New Non-Jumbo borrowers	0.25	-	-3.374***	0.168**	-	-2.003***
	(1.11)	-	(6.31)	(2.08)	-	(4.30)
Share of New Non-Jumbo borrowers	-0.209**	-	0.845**	-0.243***	-	0.401
* Saiz Elasticity of housing supply	(2.02)	-	(2.55)	(2.77)	-	(1.22)
Share Near the Jumbo Cutoff * Change in Cutoff	-	4.687***	22.91***	-	1.835**	5.376**
	-	-3.967	(7.48)	-	(1.97)	(2.62)
Share Near the Jumbo Cutoff * Change in Cutoff	-	-2.013**	-6.594***	-	-1.032***	-3.907*
* Saiz Elasticity of housing supply	-	(2.05)	(3.46)	-	(2.73)	(1.84)
Saiz Elasticity of housing supply	-0.00447***	-0.00342***	-0.00225***	-	-	-
	(4.09)	(3.47)	(2.64)	-	-	-
Time fixed effects	yes	yes	yes	yes	yes	yes
Industry structure	yes	yes	yes	yes	yes	yes
Banking Sector Controls	yes	yes	yes	yes	yes	yes
CBSA dummmies	no	no	no	yes	yes	yes
Observations	2,783	2,783	2,783	2,783	2,783	2,783
R-squared	0.316	0.322	0.347	0.524	0.516	0.525

*** p<0.01, ** p<0.05, * p<0.1

Table 6: IV Regressions relating Local Economic Growth to Housing Price Growth

This table reports IV regressions of economic growth on housing price growth by CBSA-Year; first stage results appear in Table 5. All regressions include time and CBSA fixed effects, along with measures of industry structure and the health of the local banking system. Standard errors are clustered by CBSA.

	Personal Income Growth		*Total Employment Growth*		*Employment Growth w/o Construction or Finance*		*GDP Growth*	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
House-Price Growth	0.186***	0.137***	0.222***	0.209***	0.168***	0.152***	0.259***	0.245***
	(4.25)	(3.52)	(5.83)	(5.76)	(5.12)	(4.77)	(4.66)	(4.39)
Lagged Dependent variable	-	(0.00)	-	-0.121**	-	-0.159***	-	0.0784*
	-	(0.05)	-	(2.53)	-	(2.92)	-	(1.90)
Time fixed effects	yes	yes	yes	yes	yes	yes	yes	yes
Industry structure	yes	yes	yes	yes	yes	yes	yes	yes
Banking Sector Controls	yes	yes	yes	yes	yes	yes	yes	yes
CBSA dummmies	yes	yes	yes	yes	yes	yes	yes	yes
Observations	2,783	2,783	2,783	2,783	2,783	2,783	2,783	2,783
R-squared	0.547	0.553	0.426	0.44	0.45	0.467	0.335	0.342

*** p<0.01, ** p<0.05, * p<0.1

Table 7: IV Regressions relating Local Economic Growth to Housing Price Growth , with Financial Integration Interaction

This table reports IV regressions of economic growth on housing price growth, financial integration (In CBSA ratio) and their interaction, by CBSA-Year. All three of these are treated as endogenous variables, with instruments from Table 5 plus the branching restrictions index. All regressions include time and CBSA fixed effects, along with measures of industry structure and the health of the local banking system. Standard errors are clustered by CBSA.

	Personal Income Growth	*Total Employment Growth*	*Employment growth w/o Construction or Finance*	*GDP Growth*
	(1)	(2)	(3)	(4)
House-Price Growth	-0.74	-1.10	-0.82	-0.70
	(0.59)	(0.44)	(0.65)	(0.35)
House-Price Growth *In CBSA Ratio	1.014*	1.426**	1.055*	1.044*
	(1.75)	(2.12)	(1.77)	(1.69)
In CBSA Ratio	0.06	0.13	0.157*	0.212*
	(0.99)	(1.53)	(1.75)	(1.76)
Time fixed effects	yes	yes	yes	yes
Industry structure	yes	yes	yes	yes
Banking Sector Controls	yes	yes	yes	yes
CBSA dummmies	yes	yes	yes	yes
Ch2-test for joint sig. of three endogenous variables	19.69	22.86	12.28	18.25
Observations	2,783	2,783	2,783	2,783
R-squared	0.547	0.553	0.426	0.44

*** p<0.01, ** p<0.05, * p<0.1