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Supply Restrictions, Subprime Lending and Regional US House Prices

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Abstract

With regard to the recent US house price cycle, we analyze how the interaction between housing supply restrictions, mortgage credit constraints and a price-to-price feedback loop affects house price volatility. Considering 247 Metropolitan Statistical Areas, we estimate a simultaneous boom-bust system for house prices, housing supply and subprime lending. The model accounts for regional differences in supply elasticities that are determined by local variations in topographical and regulatory supply restrictions. Our results suggest that tighter supply restrictions lead to both a larger house price boom and bust, and that this is due to supply restricted areas being significantly more exposed to a financial accelerator effect and a price-to-price expectation mechanism. We further find that the presence of endogenous price acceleration mechanisms contribute to dilute the positive relationship between the total quantity response and the supply elasticity during a housing boom.

JEL-Code: G210, R140, R210, R310.

Keywords: regional boom-bust cycles, housing supply restrictions, subprime lending, financial accelerator.

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1 Introduction

In most industrialized countries, the past decades have demonstrated a crucial role of national house price cycles in transmitting shocks to the real economy (Ferreira et al., 2010; Levitin and Wachter, 2013). However, there are large discrepancies in house price dynamics also within a particular country, and national house price cycles are often driven by developments in certain regional markets (Glaeser et al., 2008; Capozza et al., 2004; Malpezzi and Wachter, 2005). For instance, while house prices increased by more than 160 percent in some coastal areas of Florida and California from 2000 to 2006, they increased by less than 20 percent in inland open space areas of the Midwest. Against this background, we consider 247 heterogeneous US housing markets and analyze whether differences in supply restrictions, subprime lending and price acceleration mechanisms can explain the diverse price patterns observed throughout the 2000s.

A branch of the literature attributes the regional variations to heterogeneous supply side restrictions (see e.g. Malpezzi (1996), Green et al. (2005), Gyourko et al. (2008), Saiz (2010) and Glaeser (2009)). In several areas located in Florida and California, housing construction is geographically restricted by the coast line or mountains etc., while inland areas face less such restrictions. Further, some local governments try to influence building activity through their regulatory framework. Against this background, Glaeser et al. (2008) develop a theoretical model to demonstrate how differences in supply side restrictions are expected to affect price and quantity dynamics through a boom-bust cycle. Their model offers two main predictions. First, during a demand-driven housing boom, supply restricted areas primarily react by increasing house prices, while unrestricted areas absorb most of the shock through higher construction activity. Secondly, assuming supply is rigid downwards, a corresponding reduction in demand during the subsequent bust period will have a negative but equally sized impact on house prices in supply-restricted and - unrestricted areas.

Using MSA level data for the 1982-1996 US house price cycle, Glaeser et al. (2008) find that more supply-inelastic areas witnessed greater house price booms, while they do not find any robust relation between the drop in house prices during the bust and supply elasticity. This is in contrast with the results in Huang and Tang (2012), who show that house prices dropped significantly more in supply-inelastic areas during the bust period of the late 2000s.

The conflicting results in the literature might indicate that other price-stimulating mechanisms gained importance during the house price cycle of the 2000s. If price increases lead to expectations of further price increases, or to a relaxation of credit constraints, this can have a strong amplifying effect on demand (Glaeser et al., 2008; Bernanke and Gertler, 1989; Kiyotaki and Moore, 1997; Aoki et al., 2004; Iacoviello, 2005). In this paper, we demonstrate how the inclusion of such acceleration mechanisms in a model similar to Glaeser et al. (2008) change the predictions considerably. First, more supply-restricted areas are expected to experience an even stronger house price boom. Second, the stronger price increase has a positive effect on supply, diminishing the difference in the quantity response between restricted and unrestricted areas. Since both forces have a negative impact on house prices during a bust, this provides a plausible explanation as to why supply-restricted areas are observed to have experienced larger price drops during the recent housing bust.

To analyze the empirical relevance of these theoretical conjectures, we consider a simultaneous equation system for the 2000-2006 boom period. The system includes a price, a quantity and a credit relationship. The financial accelerator is captured by an endogenous feedback effect between house prices and subprime lending. We analyze how the effects of the financial accelerator deviate across areas depending on the supply elasticity. The latter is accounted for by area-specific supply elasticities depending on both topographical and regulatory supply restrictions. We also explore the relevance of a price-to-price feedback loop, by assuming that price expectations are formed adaptively, which is motivated by results in Abraham and Hendershott (1996), Shiller (2008), Case et al. (2012) and Jurgilas and Lansing (2013).

The contributions of our econometric analysis are twofold. First, our model has the advantage that it allows us to study not only the price dynamics, but also the quantity and subprime responses to, e.g., a positive housing demand shock. Second, it allows us to identify how the financial accelerator and adaptive expectations affect regional house price dynamics, and how the price, quantity and subprime responses depend on supply side restrictions.

Our econometric results confirm the main conjectures of the theoretical model. First, we find that there was both a significantly stronger financial accelerator and price-to-price feedback loop in more supply-restricted areas during the recent boom period. Second, even though these areas experience a relatively low quantity response for a given price increase, the stronger endogenous price acceleration contributes to a large increase in construction activity. In fact, our results suggest that the acceleration mechanisms contribute to diluting the relation between supply restrictions and the total quantity increase. Since housing supply is rigid downwards, this also offers a sensible explanation as to why more restricted areas were hit harder during the recent housing bust (Huang and Tang, 2012).

The financial and expectations accelerators are found to reinforce each other, with the expectations accelerator working through the financial accelerator. The financial innovations of the late 1990s - especially the introduction of subprime mortgages - may have strengthened these acceleration effects through the recent house price cycle, which provides an explanation to the diverging results found in Glaeser et al. (2008) and Huang and Tang (2012). Our results also suggest that regulatory restrictions are more important than geographical restrictions in explaining regional housing market differences. This implies that when deciding to regulate housing supply, policy makers should bear in mind how this – in combination with geographical restrictions and deregulation of financial markets – may affect the dynamics of the housing market through a boom-bust cycle.

Controlling for, among others, differences in income levels, population density, poverty rates and foreclosure laws does not materially change our results.

The paper proceeds as follows. The next section provides a theoretical motivation for the empirical analysis. In Section 3, we present our econometric models, the empirical hypotheses and describe the data that are utilized in the econometric analysis. Section 4 presents and discusses the empirical results. In Section 5, we explore the robustness of our findings, while the final section concludes the paper.

2   Theoretical motivation

2.1   A baseline supply-demand framework

Following Glaeser et al. (2008), we consider an economy consisting of several heterogeneous housing markets with different supply elasticities. Assuming that all areas initially are hit by a positive and similar sized exogenous demand shock, we analyze how the boom-bust cycle dynamics depend on supply elasticities.

In each period, the law of motion of capital accumulation for area i is given as:[1]

$$H_{i,t} \;=\; H_{i,t-1} + I_{i,t} \tag{1}$$

where $H_{i,t}$ is the housing stock at time t, while $I_{i,t}$ represents new investments in housing capital. We assume that investments are determined according to Tobin's Q theory (Tobin, 1969), i.e. new construction projects are initiated as long as the market price, $PH_{i,t}$, exceeds the marginal cost of construction, $MC_{i,t}$.

When heterogeneous areas of different sizes are considered, the number of new construction projects initiated in each period will depend on the size of the market in question. To take account of this, we assume that the marginal cost of investments is inversely proportional to the existing housing stock, i.e. there is a larger construction capacity in bigger markets. The marginal cost function for area i takes the following form:

$$MC_{i,t}(I_{i,t}) \;=\; C_{0,i}\left(I_{i,t}/H_{i,t-1} + 1\right)^{1/\varphi_i} \qquad , \varphi_i > 0 \;\; \forall \, i$$

where φ_i is the time invariant area specific supply elasticity, while $C_{0,i}$ measures fixed costs of housing construction. Setting the price equal to the marginal cost, we get the following investment function:

$$I_{i,t} \;=\; H_{i,t-1} \cdot \max\left\{0, \; \left(\frac{PH_{i,t}}{C_{0,i}}\right)^{\varphi_i} - 1\right\} \tag{2}$$

Given a non-zero supply elasticity, it follows from (2) that there will be positive investments *if and only if* prices exceed the fixed costs of construction. From (1) and (2), we find that a log transformation (lower-case letters) of the supply equation yields:[2]

$$h_{i,t} \;=\; h_{i,t-1} + \max\left\{0, \; \varphi_i\left(ph_{i,t} - c_{0,i}\right)\right\} \tag{3}$$

It follows that the log supply curve will be piecewise linear and kinked; only if the price exceeds the fixed cost of construction will supply increase as a function of the supply elasticity, φ_i, and the price-to-cost ratio (Tobin's Q). Hence, supply is assumed completely rigid downwards, motivated by the fact that houses usually neither are demolished nor dismantled (see also the discussion in Glaeser and Gyourko (2005)).

We follow custom when it comes to the modelling of the demand side. For each area, it is assumed that demand is determined in accordance with the life-cycle model of housing

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[1]We abstract from depreciation of the existing stock. Since we restrict our analysis to the short and medium run (the course of a boom-bust cycle), the depreciation will be minor and almost equal across areas.

[2]This is seen by rewriting (1) using (2); $H_{i,t} = H_{i,t-1} \cdot \max\left\{1, \; \left(\frac{PH_{i,t}}{C_{0,i}}\right)^{\varphi_i}\right\}$ and then taking logs.

(see e.g. Meen (1990, 2001) and Muellbauer and Murphy (1997), and the references therein). For area i, a logarithmic representation of the inverted demand function is given as:

$$ph_{i,t} = v_{0,i,t} + v_1 h_{i,t} \qquad , v_1 < 0 \qquad (4)$$

where $v_{0,i,t}$ is a vector product of demand shifters, such as income, the user cost of housing, and – important to the focus of this paper – credit constraints and expectations about future price changes. The parameter v_1 measures the price elasticity of an increase in the number of houses (the inverse demand elasticity).

Figure 1 illustrates the predictions of the model by showing supply-demand diagrams for a market where supply is elastic and one where it is inelastic. Initially, at time $t-1$, both markets are assumed to be in equilibrium (point A), i.e. $ph_{i,t-1} = c_{0,i}$, $I_{i,t-1} = 0$ and $H_{i,t-1} = H_{i,t-2}$. Then, in period t, both markets are hit by a positive and temporary demand shock, $\tilde{v}_{0,i,t} \in v_{0,i,t}$ increases, which triggers a one period boom. In the diagrams, this is illustrated by shifting the demand curve from D_{t-1} to D_t, and the new equilibrium is at point B. As seen, the demand shock primarily leads to quantity adjustments in the supply elastic market, and higher prices in the inelastic market.[3]

Figure 1: Boom-bust cycles in supply-elastic vs. - inelastic markets with exogenous credit provision
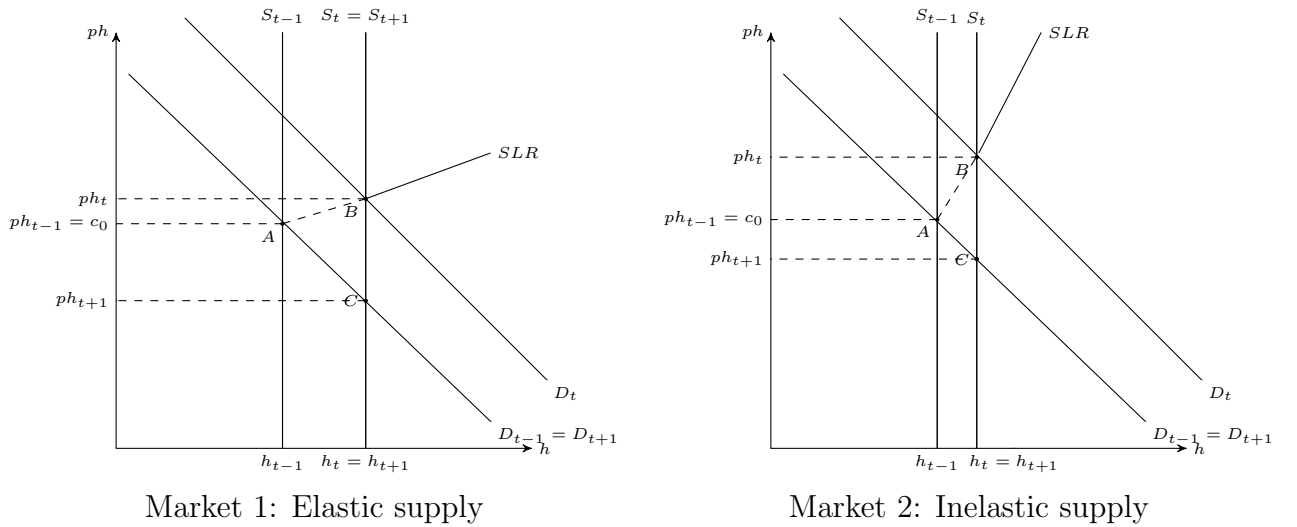
Market 1: Elastic supply   Market 2: Inelastic supply

Note: D_{t-1} is the original demand curve (which is also the demand curve in the bust period, $D_{t+1} = D_{t-1}$), while D_t is the demand curve after the positive demand shock. S_{t-1} is the original short-run supply curve and $S_t = S_{t+1}$ is the short-run supply curve after the shock materializes. The long-run supply curve is given by SLR.

In period $t+1$, demand returns to its initial level (D_t to $D_{t+1} = D_{t-1}$), which sets off a bust that also lasts for one period – the movement from point B to point C in Figure 1. Given the assumption that both markets initially are in equilibrium ($ph_{i,t-1} = c_{0,i}$), the

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[3]From the reduced form solutions of the model, the analytical expressions for the boom period responses in prices and quantity are given by: $\frac{\partial ph_{i,t}}{\partial \tilde{v}_{0,i,t}} = \frac{1}{1-v_1\varphi_i}$ and $\frac{\partial h_{i,t}}{\partial \tilde{v}_{0,i,t}} = \frac{\varphi_i}{1-v_1\varphi_i}$. The bust period price response is given by $\frac{\partial ph_{i,t+1}}{\partial \tilde{v}_{0,i,t}} = -1$. See Appendix A for details.

price will be lower than the fixed cost of construction for any value of φ_i during the bust period. It then follows from (2) that investments drop to zero, and that the price drop will be independent of the supply elasticity (see also footnote 3). Put differently, at the peak of the boom (point B in Figure 1), the *quantity overhang* is greater in the supply elastic area, whereas the *price overhang* is greater in the supply inelastic area. When demand returns to its original level during the bust, it is seen that the vertical distance from point B to C is the same in both markets, i.e. the price and quantity overhangs are equally important for the size of the bust period price drop.

The main theoretical conjectures of the model are summarized in the following proposition:

Proposition 2.1. *In a housing demand-supply model with* **exogenous** *credit provision, supply-inelastic areas will experience a greater price increase following a positive demand shock, while the quantity reaction will be greater in supply-elastic areas. When the demand shock is reversed, the price drop will be unrelated to the supply elasticity.*

Proof: See Appendix A

2.2   An extended financial accelerator model

There might be several reasons why the price and quantity dynamics through a boom-bust cycle differ from the predictions of the framework outlined in the previous section. Glaeser et al. (2008) discuss the case when price expectations are formed adaptively and show that this will generate a price-to-price feedback loop resulting in more volatile price dynamics, especially in supply-inelastic areas. In this section, we will argue that similar effects results from the existence of a financial accelerator mechanism (see e.g. Kiyotaki and Moore (1997), Bernanke and Gertler (1989), Bernanke et al. (1999), Aoki et al. (2004) and Iacoviello (2005)).

When house prices increase, households have more collateral available to pledge and, hence, banks' willingness and/or ability to lend increases. This implies that households are able to bid up prices further, possibly initiating a credit-house price spiral. We shall distinguish between lending practices in periods of non-increasing and strictly increasing house prices. In the latter case, we follow custom and assume that agents in the economy are faced with a collateral constraint in the mould of Kiyotaki and Moore (1997). In the bust period, when prices decrease, we assume that the supply of credit is fixed at some level κ_0. We then have:

$$b_{i,t} \quad \leq \quad \begin{cases} \kappa_0 + \kappa_1 ph_{i,t} & \text{, for } ph_{i,t} > ph_{i,t-1} \\ \kappa_0 & \text{, for } ph_{i,t} \leq ph_{i,t-1} \end{cases} \tag{5}$$

where $b_{i,t}$ is the log of the total amount of credit extended in area i, which during periods of strictly increasing house prices depends on house prices through the parameter κ_1.[4] We shall assume that the credit constraint is binding, and that credit is an important demand component, contained in the term $v_{0,i,t}$ in (4).[5] Thus, $v_{0,i,t}$ can be split into two

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[4]Note that by setting $\kappa_1 = 0$, we are back to the baseline model presented in the previous subsection.

[5]In practice, some people are always strictly credit constrained and unable to enter the housing market. Others will never be credit constrained, and some are in between. The composition of what people are credit-constrained or not may change over time, and for our analysis it is only important that a critical mass is credit-constrained to at least some extent, which seems to be an innocuous assumption.

components: $v_{0,i,t} = \hat{v}_{0,i,t} + \eta b_{i,t}$, where η captures the impact of credit on the demand for housing, while $\hat{v}_{0,i,t}$ measures other demand components. Solving for $b_{i,t}$ in (4), we can study how the inclusion of a financial accelerator affects the house price and quantity responses to exogenous demand shocks. In particular, a quasi-reduced form expression for housing demand is obtained by substituting (5) into (4).

$$ph_{i,t} = \begin{cases} \frac{1}{1-\eta\kappa_1}\left(\tilde{v}_{0,i,t} + \eta\kappa_0 + v_1 h_{i,t}\right) & \text{, for } ph_{i,t} > ph_{i,t-1} \\ \tilde{v}_{0,i,t} + \kappa_0 + v_1 h_{i,t} & \text{, for } ph_{i,t} \leq ph_{i,t-1} \end{cases} \tag{6}$$

Figure 2 shows similar diagrams as in the previous section, where the demand curve is now based on (6). Consider again a two period boom-bust cycle scenario, where each market is hit by a temporary and positive demand shock in period t; $\tilde{v}_{0,i,t} \in \hat{v}_{0,i,t} \subset v_{0,i,t}$ increases. In the figure, this is illustrated by the shift in the demand curve from D_{t-1} to D_t. In addition to the mechanisms described in relation to the baseline model, the increase in house prices cause banks to extend more credit, which again spurs house prices and construction activity. The importance of the feedback mechanism between house prices and credit is captured by the magnitude of $\eta\kappa_1$, i.e. the steepness of the demand curve during periods of increasing house prices. The higher $\eta\kappa_1$ is, the more prices will accelerate during the boom, and the acceleration is greater when supply is inelastic. Depending on the numerical size of $\eta\kappa_1$, the quantity increase may be independent of the supply elasticity.[6]

Turning to the bust period, we assume that demand returns to its pre-boom level. Given that house prices are falling, the amount of credit will be fixed, such that the demand curve gets a new kink. Since housing supply is also fixed at the boom period level due to the durability of housing, the model predicts that the price drop is significantly greater in supply inelastic areas. Compared with the baseline model, the additional credit driven price and quantity overhangs are relatively larger in supply inelastic areas, and they both have a negative effect on prices during the bust (see also footnote 6).

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[6]From the reduced form solutions of the model, the analytical expressions for the boom period responses in prices and quantity are given by: $\frac{\partial ph_{i,t}}{\partial \tilde{v}_{0,i,t}} = \frac{1}{1-v_1\varphi_i-\eta\kappa_1}$ and $\frac{\partial h_{i,t}}{\partial \tilde{v}_{0,i,t}} = \frac{\varphi_i}{1-v_1\varphi_i-\eta\kappa_1}$. The bust period price response is given by $\frac{\partial ph_{i,t+1}}{\partial \tilde{v}_{0,i,t}} = \frac{1-v_1\varphi_i}{1-v_1\varphi_i-\eta\kappa_1}$, see Appendix A for details. If $\eta\kappa_1 > 1$, the quantity increase will be greater in supply inelastic markets.

Figure 2: Boom-bust cycles in supply-elastic vs. - inelastic markets with endogenous credit provision
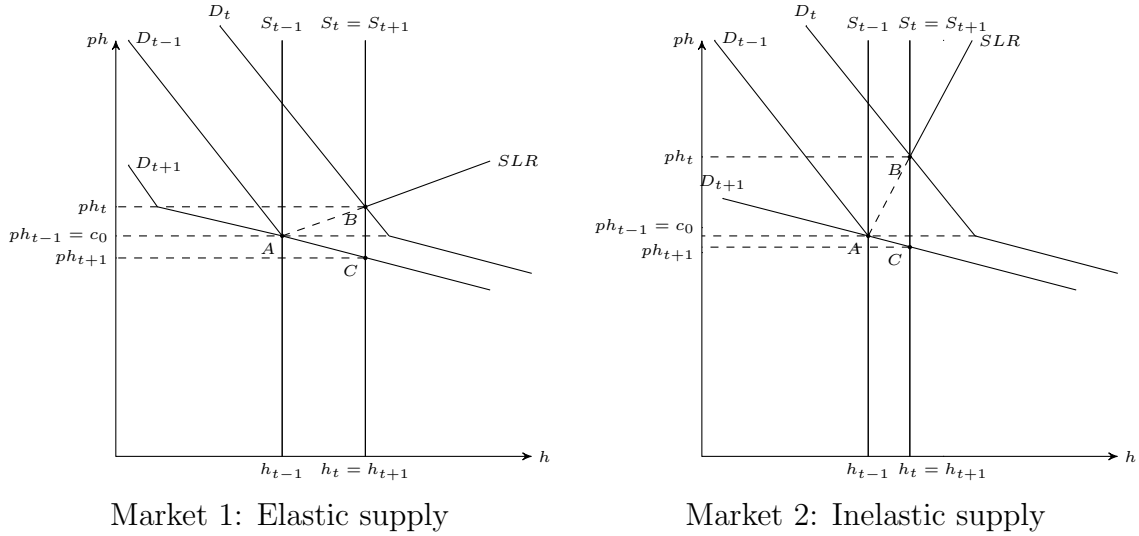
Market 1: Elastic supply   Market 2: Inelastic supply

Note: D_{t-1} is the original demand curve (which is also the demand curve in the bust period, $D_{t+1} = D_{t-1}$), while D_t is the demand curve after the positive demand shock. S_{t-1} is the original short-run supply curve and $S_t = S_{t+1}$ is the short-run supply curve after the shock materializes. The long-run supply curve is given by SLR.

The main results for this extended financial accelerator model are summarized in the following proposition:

Proposition 2.2. *In a housing demand-supply model with **endogenous** credit provision, supply-inelastic areas will experience a greater price and credit increase following a positive demand shock, while the quantity reaction may be either higher or lower than that of supply-elastic areas. When the demand shock is reversed, the price drop will be greater in supply-inelastic areas.*

Proof: See Appendix A

In terms of implied price dynamics, our theoretical results are similar to Glaeser et al. (2008), who consider a price-to-price feedback loop. Further, a combination of the two mechanisms could also be relevant, which is in line with Brueckner et al. (2012), who argue that the explosion in subprime lending was both a cause and a consequence of bubble-like behavior in the US housing market in the 2000s. In particular, they argue that adaptive house price expectations spurred subprime lending by easing default concerns. As a result of this, housing demand increased, which lead to higher prices etc. In a dynamic setting, this would make the price-accelerating mechanisms even stronger than what is found in this section, and contribute to further diminish the differences in the quantity response between supply-elastic and - inelastic areas. In the empirical analysis, we investigate the relevance of adaptive price expectations by including lagged house price appreciation as an explanatory variable in our econometric model.

3 Econometric model and data

3.1 Econometric approach

Our econometric framework departs from the supply-demand relations in (3) and (4). Considering this model in differences, we arrive at our baseline simultaneous demand-supply system for the boom period:

$$\Delta ph_i^{Boom} = \alpha_1 + \beta_{1,\Delta h}\Delta h_i^{Boom} + \boldsymbol{\beta}_{1,\boldsymbol{x}}'\boldsymbol{x}_i^{Boom} + \varepsilon_{\Delta ph,i} \tag{7}$$

$$\Delta h_i^{Boom} = \alpha_2 + \left(\beta_{2,\Delta ph} + \boldsymbol{\beta}_{2,\Delta ph\times\boldsymbol{Reg}}' \times \boldsymbol{Reg}_i\right)\Delta ph_i^{Boom} + \beta_{2,\boldsymbol{z}}\boldsymbol{z}_i + \varepsilon_{\Delta h,i} \tag{8}$$

where Δph_i^{Boom} and Δh_i^{Boom} represent the boom period growth in house prices and quantity for area i, respectively. $\boldsymbol{Reg}_i$ is a vector of supply restriction measures, affecting the area specific supply elasticities. The empirical counterpart of φ_i in (3) is given by $\beta_{2,\Delta ph} + \boldsymbol{\beta}_{2,\Delta ph\times\boldsymbol{Reg}}' \times \boldsymbol{Reg}_i$, which henceforth will be referred to as the implied supply elasticity. The demand shifters, $v_{0,i,t}$ in (4), are collected in the vector product $\beta_{1,\boldsymbol{x}}\boldsymbol{x}_i$, with $\boldsymbol{x}_i = \left(\Delta y_i^{Boom}, \Delta sp_i^{Boom}\right)'$, where Δy_i^{Boom} measures growth in income in area i and Δsp_i^{Boom} is the log cumulative increase in subprime originations per capita.[7] Finally, $\boldsymbol{z}_i = \left(\Delta cc_i^{Boom}, \boldsymbol{Reg}_i'\right)'$, where Δcc_i^{Boom} is growth in construction wages. These include both fixed and variable costs of construction, which we disregarded in the theoretical exposition.

Later, we shall extend this baseline model to allow for a financial accelerator effect. With reference to equation (5), we assume the following relationship for subprime lending:

$$\Delta sp_i^{Boom} = \alpha_3 + \beta_{3,\Delta ph}\Delta ph_i^{Boom} + \boldsymbol{\beta}_{\boldsymbol{y}}'\boldsymbol{y}_i^{Boom} + \varepsilon_{\Delta sp,i} \tag{9}$$

where $\boldsymbol{y}_i^{Boom} = \left(\Delta y_i^{Boom}, \text{Denial share}_{1996}, \text{LTI}_{1996}\right)$. We follow Mian and Sufi (2009) and Huang and Tang (2012) and use the loan denial rates in 1996 (Denial share$_{1996}$) as an instrument for subprime lending. Mian and Sufi (2009) argue that areas with high rejection rates before the boom period were more likely to be exposed to subprime lending at a later stage, since the pool of borrowers falling into this category was larger in these areas. The second instrument, the average loan-to-income ratio (LTI$_{1996}$), has been considered by Wheaton and Nechayev (2008) as a proxy for looser lending standards. The specification in (9) implies that house prices can influence subprime lending, which allows for the possibility of a financial accelerator.

As alluded to in the previous section, we will also extend our model to allow for a price-to-price feedback loop. Motivated by findings in Abraham and Hendershott (1996), Shiller (2008), Case et al. (2012) and Jurgilas and Lansing (2013), we augment our econometric system by lagged house price changes, $\Delta ph_i^{Pre-boom}$, in order to capture an extrapolative expectation formation. Expectations may affect housing demand both *directly*, as argued in e.g. Glaeser et al. (2008) and Shiller (2008), and *indirectly* through more lending (see e.g. Brueckner et al. (2012)). Housing supply may also be affected, since expectations about future prices may be important in a tedious construction and planning process. For this reason, we shall allow lagged price changes to influence both demand, supply and subprime lending.

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[7]Since the interest rate is almost equal across areas, we abstract from the user cost component.

Our econometric framework is complicated by the region-specific elasticities, modelled in (8) as interaction terms between house price changes and the supply restriction indexes. However, following the argument in Wooldridge (2010, Ch. 9), this non-linearity does not change the identification requirements of the model, which are indeed satisfied in all cases.[8] The systems are estimated by full information maximum likelihood (FIML), assuming that the disturbances follow a joint normal distribution.

The baseline boom system, as given by (7) and (8), is related, and to some extent comparable, to the reduced form specifications considered in Glaeser et al. (2008) and Huang and Tang (2012). However our model is structural, allowing us to study the impact on both prices and quantity in light of exogenous shocks. In addition, our modelling strategy gives rise to a non-linear interaction between price and quantity responses and the region-specific supply restrictions, as described above. We will, however, compare the main predictions and the qualitative results of the models.

3.2 Data definitions

Our data set originally covers 247 US Metropolitan Statistical Areas (MSAs).[9] However, we have excluded five areas from our sample. In these areas, housing market dynamics in the 2000s were dominated by extreme exogenous shocks, which are unrelated to the questions we raise in this paper.[10] We follow Huang and Tang (2012) and define the boom as the period from 2000 to 2006, while the bust runs from 2006 to 2010.

A large number of data sources have been utilized to construct our data set. Our data on lending conditions are constructed based on the Home Mortgage Disclosure Act (HMDA) loan application registry (LAR) data (see Appendix B for details).[11] These data cover information on loan applications for about 92 percent of the US population, including the number of applications, the income of the applicant, loan amount, whether the loan was denied or originated, and whether the financial institution extending the loan was engaged in subprime lending.[12] We use the data at the loan applicant level to construct the log cumulative number of subprime originations per capita (measured per 100 people) at the MSA level during the boom period. In addition, the data are used to construct the 1996 denial share and LTI ratio.

Data on household income and the housing stock have been collected from Moody's Analytics, while the house price data are from the Federal Housing Finance Agency

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[8]In the baseline model (see (7) and (8)), the subprime measure and the income variable only enter the price equation, while construction wages and the non-interacted restriction indexes only enter the supply equation. This means that we have an overidentified system with two testable exclusion restrictions.

[9]We use the 2004 MSA definitions of the Census Bureau. Details on the MSAs included in our analysis are available upon request.

[10]Four MSAs situated in Louisiana experienced a large negative shock to housing supply through Hurricane Katrina and the subsequent floods in late August 2005. These are New Orleans-Metairie-Kenner (LA), Lake Charles (LA), Alexandria (LA) and Monroe (LA). Such episodes are hard to reconcile with any plausible economic interpretation, and must be interpreted as extraordinary circumstances. We also excluded Barnstable Town (MA), due to extreme degrees of political and geographical supply restrictions.

[11]For a summary of the opportunities and limitations of the HMDA data, see the discussion in Avery et al. (2007).

[12]To determine this, we had to match the HMDA data with the subprime list provided by the Department of Housing and Urban Development (HUD).

(FHFA) and construction wages are taken from the St. Louis Fed's database FRED.[13] All variables are measured in nominal terms.

To control for regional differences in supply restrictions, we consider two indexes. First, we use the Wharton Regulatory Land Use Index (WRLURI) from Gyourko et al. (2008). WRLURI measures MSA level regulatory supply restrictions, including complications related to getting a building permit etc.[14] Second, to measure topographical supply restrictions, we consider the UNAVAL index by Saiz (2010). This index measures MSA-level geographical land availability constraints.[15] An advantage of the index developed by Saiz (2010) is that nature given supply restrictions are exogenous to housing market conditions. This is not necessarily the case for regulatory supply restrictions, which could bias our results.[16] It should, however, be noted that the other coefficients in our model are relatively invariant to leaving out this index.

As noted by Glaeser et al. (2008), the two supply restriction indexes are positively correlated. Instead of leaving out one index – the approach pursued in Glaeser et al. – we assume that UNAVAL is exogenous and use this index as is, while WRLURI is adjusted for the possible influence of UNAVAL.[17] In order to make the estimated effect of the two indexes comparable, we normalize WRLURI to range between 0 and 1 in the original sample. The adjusted index is labelled WRLURI(a) and is orthogonal to UNAVAL. We should be able to interpret UNAVAL as an exogenous effect of physical supply restrictions. However, some of the observed effect of UNAVAL might be caused by more geographically constrained areas having more regulations on building permits etc., possibly to preserve nature.

3.3 Descriptive statistics

There are substantial differences across the MSAs covered by our sample. During the 2000–2006 boom period, nominal house price growth ranges from more than 160 percent in Naples-Marco Island (FL) to a little less than 10 percent in Lafayette (IN). Similar variations are seen in the boom period change in quantity and subprime mortgage extensions, as well as in the bust period price changes.

There are also large variations in the supply restriction indexes. The geographical land restriction measure (UNAVAL) indicates that only 0.05 percent of the land is rendered

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[13]This variable is (for natural reasons) highly correlated with income growth, so we include the orthogonalized measure, i.e the residual from the following regression: $\Delta cc_i = \beta_0 + \beta_1 \Delta y_i + \varepsilon_i$.

[14]This index is originally based on 11 subindexes measuring different types of complications and regulations in the process of getting a building permit. WRLURI is available at a town (or city) level, which we have aggregated to the MSA level using the sample probability weights of Gyourko et al. (2008).

[15]Saiz (2010) uses GIS and satellite information to calculate the share of land in a 50 kilometre radius of the MSA main city centres that is covered by water, or where the land has a slope exceeding 15 degrees. These areas are seen as severely constrained for residential construction.

[16]It is not clear in which direction the bias would go in terms of the effects of the regulatory supply restrictions: If house prices increase, building activity will increase as well. To constrain the high building activity, local governments might respond by enforcing more restrictions. On the other hand, booming house prices are often accompanied by increasing economic activity, job creation, population growth etc. In order to dampen the pressure on house prices, or to provide homes for an increasing population, governments might relax regulations on construction activity.

[17]We use the residuals from the following specification to measure the part of WRLURI that is not explained by UNAVAL: $WRLURI_i = \beta_0 + \beta_1 UNAVAL_i + \varepsilon_i$.
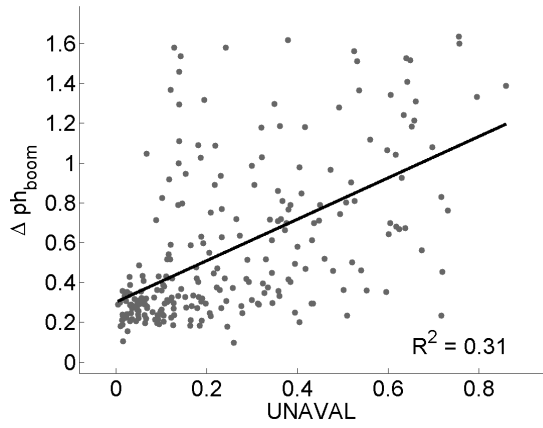
undevelopable in Lubbock (TX), while as much as 86 percent of the land is considered undevelopable in Santa Barbara-Santa Maria-Goleta (CA). Regarding our measure of regulatory supply side restrictions (WRLURI(a)), Glens Falls (NY) is the least restricted area. Despite the high geographical supply restrictions in the area, it has a low degree of political involvement in the development process, low requirements for developers and a fast building permit application process (WRLURI(a) = 0.01). At the other extreme, even after controlling for a high degree of geographical supply restrictions, Boulder (CO) has a very high level of political involvement in the urban development process and a long and complex building application process etc. (WRLURI(a) = 0.74).[18]

In order to get a first-hand idea of the correlation among the variables, Panel (a)–Panel (d) in Figure 3 plots the growth in house prices and quantity over the 2000-2006 period against the two supply restriction indexes. It is clear that more supply restricted areas – both geographically and regulatory – experienced a greater house price boom during the 2000s. On the other hand – and this is a puzzle for the baseline theory model – there does not seem to be any systematic link between supply restrictions and the increase in quantity over the boom period. Unless supply inelastic markets were systematically hit by larger (or more) demand shocks, this is hard to reconcile with the predictions of the baseline theory model.

Turning to the bust price growth, Panel (e) and Panel (f) in Figure 3 plots the price growth during the bust against the price and quantity growth during the boom. As seen, both the price and quantity overhang are negatively correlated with price movements during the bust.

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[18]In the original sample Barnstable Town (MA) was the most regulated area (WRLURI(a) = 1), while New Orleans-Metairie-Kenner (LA) was the least regulated area (WRLURI(a) = 0).

Figure 3: Supply restrictions and price and quantity changes over the boom-bust cycle from 2000-2006

(a) Δph^{Boom} vs. UNAVAL
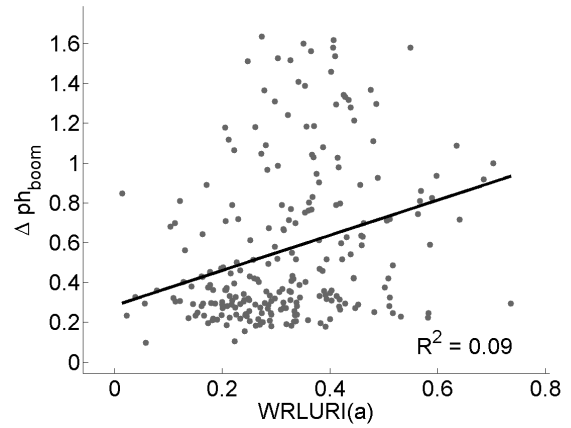
(b) Δph^{Boom} vs. WRLURI(a)

(c) Δh^{Boom} vs. UNAVAL

(d) Δh^{Boom} vs. WRLURI(a)

(e) Δph^{Bust} vs. Δph^{Boom}

(f) Δph^{Bust} vs. Δh^{Boom}

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4 Empirical results

4.1 The baseline boom period model

The results obtained when estimating the boom system in (7) and (8) are displayed in Table 1.

Table 1: The baseline supply-demand system for the boom period, 2000–2006

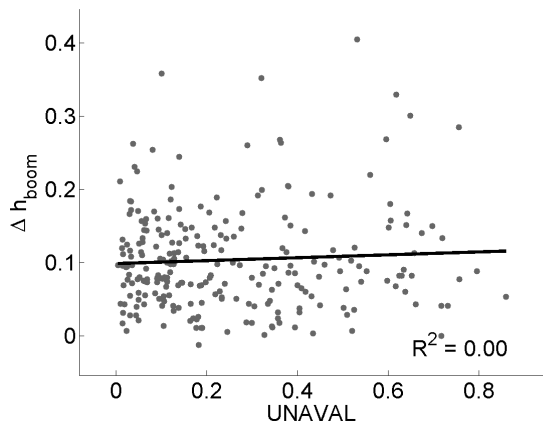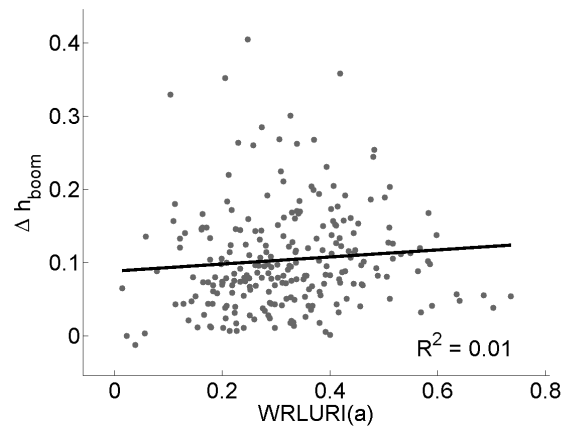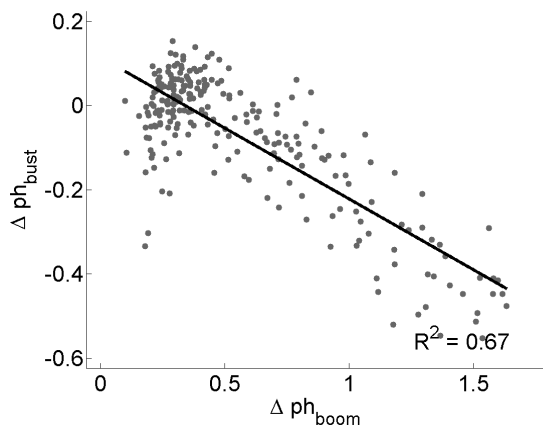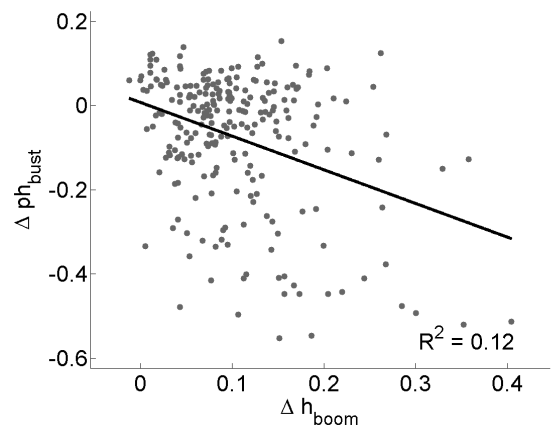| Variable	Δph_{Boom}		Δh_{Boom}	
	Coefficient	t-value	Coefficient	t-value
Constant	3.050	6.166	-0.055	1.534
Δh_{Boom}	-13.977	5.587	*	*
Δph_{Boom}	*	*	0.533	7.979
$una \times \Delta ph_{Boom}$	*	*	-0.128	2.117
$wrl \times \Delta ph_{Boom}$	*	*	-0.600	5.061
Δsp_{Boom}	0.507	6.556	*	*
Δy_{Boom}	6.434	6.419	*	*
Δcc_{Boom}	*	*	-0.202	4.838
una	*	*	-0.149	2.504
wrl	*	*	0.113	1.330
Overidentifying restriction ($\chi^2(3)$)	2.2044 [0.5311]			
Normality test	p-value= 0.0000			
Observations	242			

Note: The table reports the FIML estimates of the boom system (7)–(8). The following abbreviations apply: h is the log housing stock, ph is log house prices, sp is log cumulative subprime originations per capita, y is household income, cc is log of construction wages, una is the geographical restriction index of Saiz (2010), wrl is the regulatory index of Gyourko et al. (2008) adjusted for una and normalized to range between 0 and 1. All variables are nominal, and all variables except the subprime variable are in percentage changes. Δ is a difference operator. An asterisk indicates that the variable is not included in the equation under consideration. The reported t-values are measured in absolute terms. The normality test is the Doornik and Hansen (2008) normality test for multivariate normality.

The two equations are interpretable as a supply-demand system: changes in quantity have a significant negative impact on house prices in the first equation (the inverted demand equation), while house prices enter positively in the second equation (the supply equation). It is clear that more supply restrictions – both regulatory and topographical - significantly lower the implied supply elasticity. Furthermore, our results imply that the more restricted the supply is, the more house prices will increase for a given positive demand shock – a finding that is consistent with the reduced form results of Glaeser et al. (2008) and Huang and Tang (2012). We also find that an increase in the subprime exposure leads to a positive reaction in house prices, in line with the results of Huang and Tang (2012).

The income coefficient in the demand equation has an expected positive sign, while construction wages are found to affect housing supply negatively. Several coefficients are quite large in numerical size compared to typical estimates of key long-run elasticities found in the literature (see e.g. Meen (2001)and Girouard et al. (2006)). That said, it must be remembered that we consider a period of strong house price appreciation, and that our estimates are better regarded as medium-term elasticities. In addition, we also have to account for the reaction in housing supply to determine the total price response. Taking account of the reaction in supply, we find that an increase in income growth of

one percentage point increases house prices by 1.28 percentage points for an area with average supply elasticity, which does not seem unreasonable.[19] On the other hand, the coefficient on construction wages may seem a bit low, from the perspective of Tobin's Q theory. This may be due to the fact that construction wages is a crude proxy for actual construction costs. The coefficient on subprime lending is relatively low due to the way the variable is measured, but the cross-MSA variation in this variable is large, so differences in subprime exposure accounts for important differences across areas.

Looking at the lower part of Table 1, it is evident that the test for overidentifying restrictions suggests that the three overidentifying restrictions – that income does not enter the supply equation and that the two restriction indexes (non-interacted) only affect housing supply – are supported (p-value = 0.53). One caveat with the model specification is that there is evidence of non-normality. However, as we show in a separate section, our results are indeed robust to this.

Figure 4 demonstrates the importance of the heterogeneous supply elasticities more clearly. Based on the results in Table 1, we calculate the house price and quantity reactions to a one standard deviation increase in subprime lending. The figure shows the response functions for house prices and quantity against the implied supply elasticity, ordered from the least to the most elastic area.[20] From Panel (a), it is clear that an increase in subprime lending affects house prices positively in all areas. That said, the price response differs quite markedly across areas, with a smaller response for areas with low implied supply elasticities. In the area with the lowest implied supply elasticity (Boulder, CO), a one standard deviation increase in subprime lending per capita leads to a 28.2 percentage point increase in the cumulative growth in house prices through the boom. The same number is only 5.5 percentage points, i.e. only a fifth, in the area with the highest supply elasticity (Pine Bluff, AR).

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[19]For the area in our sample with the lowest supply elasticity, the same number is 4.25 percentage points, while it is 0.82 percentage points for the area with the highest implied elasticity.

[20]Standard errors are calculated using the delta method, see Appendix E for details.

Figure 4: Price and quantity responses for different supply elasticities for baseline model

(a) House prices
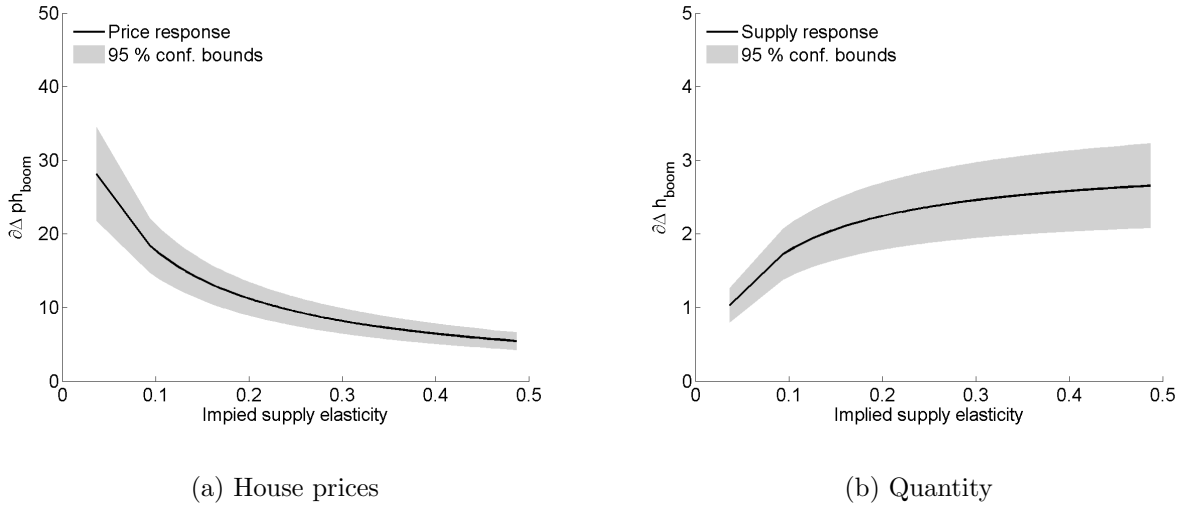
(b) Quantity

Note: This figure shows the boom period price and quantity responses to a one standard deviation increase in the cumulative change in subprime lending per capita. The calculations are based on the first derivatives of the reduced form house price and quantity equations, and the confidence bounds are calculated using the delta method. See Appendices D and E for details.

Turning to the supply side of the model, Panel (b) in Figure 4 shows the quantity response for the same shock in subprime lending. The average response is an increase in quantity of 2.4 percentage points. Hence, there is a considerably smaller change in quantity than in house prices, as would be expected. In support of the theoretical model, we find that the quantity responses are greater for areas with higher implied supply elasticities. However, the variation in the quantity responses is much smaller than the variation in the price responses. In particular – and this is a puzzle for the baseline model presented in Section 2 – there is literally no significant difference in the quantity response across a majority of the areas.

4.2 The extended financial accelerator boom period model

In this section, we explore the empirical relevance of the financial accelerator by letting subprime lending be endogenously determined in our system, as given by (9). The results of the extended boom system are reported in Table 2. It is evident that all the coefficients in the second equation (the supply equation) are close to those reported in Table 1.[21] The coefficients in the first equation (the inverted demand equation) are also similar to those of the baseline model, though treating subprime as exogenous (as in the previous section) biases the coefficient on the change in quantity downwards and the coefficient on income upwards. Turning to the final equation (the subprime equation), house prices are found to exercise a positive and highly significant effect on regional subprime extensions.

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[21]The implied elasticities of this model have a correlation of around 0.7 with the implied elasticities found by Saiz (2010), who uses a different approach to pin down MSA specific elasticities. The correlation between the elasticities implied by this extended model and those implied by the baseline model is 0.99.

Combined with the positive effect of subprime lending on house price growth, this gives rise to a financial accelerator mechanism where higher house prices increase subprime lending, and *vice versa*. Moreover, since the price response to a given shock is predicted to be greater in more supply-restricted areas, there will also be a larger credit multiplier in these areas.

Table 2: The boom period model including a financial accelerator, 2000–2006

Variable	Δph_{Boom}		Δh_{Boom}		Δsp_{Boom}	
	Coefficient	t-value	Coefficient	t-value	Coefficient	t-value
Constant	3.529	6.985	-0.058	1.585	-9.833	23.463
Δh_{Boom}	-11.376	5.373	*	*	*	*
Δph_{Boom}	*	*	0.541	7.829	0.686	3.553
$una \times \Delta ph_{Boom}$	*	*	-0.171	3.016	*	*
$wrl \times \Delta ph_{Boom}$	*	*	-0.578	4.979	*	*
Δsp_{Boom}	0.559	7.447	*	*	*	*
Δy_{Boom}	5.121	5.626	*	*	1.248	2.998
Δcc_{Boom}	*	*	-0.203	4.962	*	*
una	*	*	-0.114	1.953	*	*
wrl	*	*	0.093	1.110	*	*
Denial share$_{1996}$	*	*	*	*	0.786	1.580
LTI$_{1996}$	*	*	*	*	1.708	7.734
Overidentifying restriction ($\chi^2(4)$)	5.8496 [0.2107]					
Normality test	p-value= 0.0002					
Observations	242					

Note: The table reports the FIML estimates of the boom system (7)–(9). The following abbreviations apply: h is the log housing stock, ph is log house prices, sp is log cumulative subprime originations per capita, y is log household income, cc is log construction wages, una is the geographical restriction index of Saiz (2010), wrl is the regulatory index of Gyourko et al. (2008) adjusted for una and normalized to range between 0 and 1, LTI_{1996} is the loan-to-income ratio in 1996 and Denial share$_{1996}$ is the denial share in 1996. All variables are nominal, and all variables except the subprime variable are in percentage changes. Δ is a difference operator. An asterisk indicates that the variable is not included in the equation under consideration. The reported t-values are measured in absolute terms. The normality test is the Doornik and Hansen (2008) normality test for multivariate normality.

The current specification implies an additional four overidentifying restrictions relative to the baseline model, since both $WRLURI(a)$ and $UNAVAL$ are excluded from the subprime equation, while the LTI ratio is excluded from both the housing demand and the housing supply equations. Testing the validity of these restrictions, we find a p-value of 0.21. The denial rate seems to be a rather weak instrument, with a t-value of only 1.58, but it should be noted that we get similar results if we drop this variable from the system altogether, i.e. only use the LTI to identify the subprime equation.

To analyze the effect of the financial accelerator, Figure 5 shows the same response graphs as in the previous section, where we also report the response functions when the financial accelerator is "switched off" by counterfactually setting $\beta_{3,\Delta ph} = 0$. As is evident from inspecting Panel (a) in Figure 5, the response without the financial accelerator is closely in line with the response in Figure 4. It is seen that the financial accelerator contributes to magnify the price reaction in all areas, but more so in areas with a lower implied supply elasticity. The average price response is 12 percentage points, and the financial accelerator increases the response – relative to the baseline model – by 11 percent on average. For the area with the lowest implied supply elasticity, the response is 43.1

Figure 5: Price and quantity responses for different supply elasticities for extended model

(a) House prices   (b) Quantity
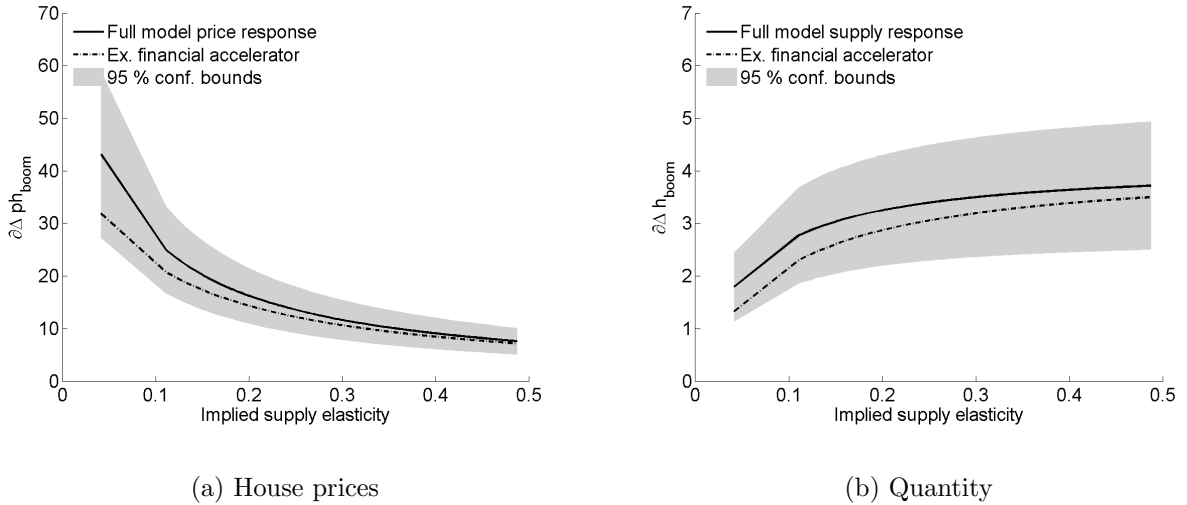
Note: This figure shows the total boom period price and quantity responses to a one standard deviation increase in the cumulative change in subprime lending per capita both with and without the financial accelerator in the model. The calculations are based on the first derivatives of the reduced form price and quantity equations, and the confidence bounds are calculated using the delta method. See Appendices D and E for details.

percentage points, and the financial accelerator increases the price response by 35 percent. For the area with the highest supply elasticity, prices increase by 7.6 percentage points, which is 6 percent higher than in the case where the financial accelerator is not accounted for.

Turning to the quantity responses, as shown in Panel (b) in Figure 5, it is evident that supply increases more in all areas when accounting for the financial accelerator. However, as expected from Proposition 2.2, we see that the quantity responses are more equal across areas. In fact, with the financial accelerator in the model, we cannot reject the hypothesis of no significant difference in the quantity reaction across a majority of the areas. This suggests that the momentum effect of the financial accelerator causes the connection between the total quantity response and the elasticity of supply to literally vanish.

The importance of credit markets in explaining regional house prices has also been addressed in other parts of the literature. One part looks at how imbalances in credit markets may generate imbalances in housing markets. Wheaton and Nechayev (2008) find that the US housing market disequilibria during the 2000s were largest in areas where the subprime mortgage market was most active – consistent with the above results. The authors are tacit about the mechanisms causing these differences, and our results suggest that one such mechanism may be the financial accelerator effect.

Another related branch of the literature is concerned with the causes of regional credit expansions. Mian and Sufi (2009) analyze regional credit market dynamics through the late 1990s and early 2000s. They do not find that large credit expansions were related to tighter supply restrictions. This leads them to reject the hypothesis of an expectations

driven credit expansion. Compared with their study, we ask whether the credit expansions were caused by more aggressive house price increases in supply inelastic areas, and not by the restrictions *per se*.

4.3   A price-to-price feedback loop

An alternative, or complementary, mechanism that can generate similar results as the financial accelerator is a price-to-price feedback loop. The existence of a price-to-price spiral rests on the premise that there is an extrapolative element in households' expectation formation, whereby increasing house prices results in expectations about higher future house prices, which spurs house price growth further, possibly through an increased credit growth. Consistent with this view, Abraham and Hendershott (1996) interpret the coefficients on lagged house price appreciation in a dynamic panel model for house prices as capturing a "bubble builder" – or a momentum – effect.

That said, the assumption that house price expectations are formed adaptively rather than rationally calls for some justification. In the literature, there seems to be strong evidence pointing in the direction that house price expectations are indeed formed in an adaptive manner (see e.g. Jurgilas and Lansing (2013) and the references therein). Shiller (2008) presents evidence from survey data from major US cities for the years 2006 and 2007 showing that individuals living in areas with a recent increase in home values expected further price increases, while the opposite was the case in areas with recent price declines. Strikingly, conducting a similar survey in the midst of the national housing bust, Case et al. (2012) find that individuals living in previously booming areas now expected a decline in house prices, and conclude that "1-year expectations are fairly well described as attenuated versions of lagged actual 1-year changes [...]". The strong correlation between survey data on house price expectations and past house price increases are further documented in Williams (2013).

While a dynamic econometric model is better suited to study the price-to-price feedback loop in detail, we will explore if there are signs that lagged house prices – interpreted as capturing adaptive price expectations – enter significantly in our boom system. In particular, we use the period from 1996 to 2000 as the "early-boom" stage, and add the house price growth over this period, $\Delta ph_{Pre-boom}$, to all equations in the boom system.[22]

Results are given in Table 3, and it is evident that lagged house prices only have a significant effect in the subprime equation. Thus, to the extent that early-boom house price developments capture an expectation channel, they seem to operate via the lending channel and not directly through housing demand or supply. Thus, an expectation of higher house prices is primarily materialized into increased demand for housing because banks loosen their lending standards and/or because households increase their demand for credit. In any case, the results indicate that higher expected prices results in more lending, which again spurs house prices through an increase in housing demand – a result that is consistent with Brueckner et al. (2012).

The coefficients in the housing supply and demand equations are robust to this model extension. This implies that the price-to-price feedback loop offers a complementary, and

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[22]Glaeser et al. (2008) use the period 1996-2006 as the latest boom period in the US housing market. Thus, the 1996–2000 period can be seen as the first stage of the boom.

Table 3: The boom period model including a financial accelerator and adaptive expectations, 2000–2006

Variable	Δph_{Boom}		Δh_{Boom}		Δsp_{Boom}	
	Coefficient	t-value	Coefficient	t-value	Coefficient	t-value
Constant	3.945	5.690	-0.047	0.912	-10.070	25.222
Δh_{Boom}	-11.770	5.655	*	*	*	*
Δph_{Boom}	*	*	0.564	7.355	0.402	2.211
$una \times \Delta ph_{Boom}$	*	*	-0.167	2.494	*	*
$wrl \times \Delta ph_{Boom}$	*	*	-0.617	4.924	*	*
$\Delta ph_{Pre-Boom}$	-0.367	0.657	-0.063	0.230	1.980	4.139
$una \times \Delta ph_{Pre-Boom}$	*	*	-0.076	0.232	*	*
$wrl \times \Delta ph_{Pre-Boom}$	*	*	0.553	0.976	*	*
Δsp_{Boom}	0.608	6.511	*	*	*	*
Δy_{Boom}	5.083	5.421	*	*	2.082	4.951
Δcc_{Boom}	*	*	-0.209	4.579	*	*
una	*	*	-0.133	1.663	*	*
wrl	*	*	0.004	0.032	*	*
Denial share$_{1996}$	*	*	*	*	0.757	1.620
LTI$_{1996}$	*	*	*	*	1.567	7.384
Normality test	p-value= 0.0005					
Observations	242					

Note: The table reports the FIML estimates of the boom system (7)–(9). The following abbreviations apply: h is the log housing stock, ph is log house prices, sp is log cumulative subprime originations per capita, y is log household income, cc is log construction wages, una is the geographical restriction index of Saiz (2010), wrl is the regulatory index of Gyourko et al. (2008) adjusted for una and normalized to range between 0 and 1, LTI_{1996} is the loan-to-income ratio in 1996 and Denial share$_{1996}$ is the denial share in 1996. All variables are nominal, and all variables except the subprime variable are in percentage changes. Δ is a difference operator. An asterisk indicates that the variable is not included in the equation under consideration. The reported t-values are measured in absolute terms. The normality test is the Doornik and Hansen (2008) normality test for multivariate normality.

not a mutually exclusive, explanation of the inter-MSA boom dynamics.[23]

4.4   The bust period

In this section, we analyze how the heterogeneous price and quantity responses during the boom may affect the size of the house price bust. Since housing is durable, we abstract from modelling the supply side in the bust period, which is also consistent with the theoretical model. Our specification for bust period house price growth is given by:

$$\Delta ph_i^{Bust} \;=\; \mu + \gamma_{\Delta ph}\Delta ph_i^{Boom} + \gamma_{\Delta h}\Delta h_i^{Boom} + \gamma_{\Delta y}\Delta y_i^{Bust} + e_i \qquad (10)$$

where Δy_i^{Bust} is the growth in disposable income during the 2006–2010 bust period. The bust equation is estimated by OLS, and results are shown in Table 4.[24]

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[23]We have also considered a quasi-reduced form representation of the system by substituting out for (9) in (7). In this specification, we find that lagged house prices enter significantly in the housing demand relation. The three-equation system does, however, suggest that this effect operates through the credit channel. Detailed results are available upon request.

[24]Introducing this equation does not cause any problems with identification, since it is recursively determined, i.e. there is no feedback from the bust price response to either the boom price or the boom supply response.

Table 4: Bust period model, 2006–2010

Variables	Coefficient	t-value
$Constant$	0.027	1.954
Δh_{boom}	-0.407	5.774
Δph_{boom}	-0.281	21.526
Δy_{bust}	0.875	11.328
Normality test	0.0000	
Obs.	242	

Note: The table reports the bust period OLS estimates of equation (10). The following abbreviations apply: h is the log housing stock, ph is log house prices, y is log household income. All variables are nominal, and all variables are in percentage changes, where Δ is a difference operator. The reported t-values are measured in absolute terms. The normality test is the Doornik and Hansen (2008) normality test for univariate normality.

As expected, the effect of both the price and the quantity overhangs are negative and highly significant. The importance of supply restrictions for the bust price response boils down to a question of how the boom period price and quantity responses depend on supply restrictions.

When accounting for the financial accelerator, we have already seen how the boom period price response is unambiguously higher in *more* supply restricted areas, while the response in quantity is practically unrelated to supply-side restrictions (see Figure 5). In a dynamic model, the price-to-price feedback loop would only strengthen this finding. Hence, in the presence of one or more price accelerating mechanisms, one would also expect the bust price response to be significantly larger in more supply restricted areas. Since subprime lending is a relatively new phenomena, this also provides an explanation of the diverging results found in Glaeser et al. (2008) and Huang and Tang (2012).

Figure 6 shows the bust period price response to the out-of-equilibrium behavior caused by the one standard deviation increase in subprime lending per capita during the boom. The response is shown both *with* and *without* the financial accelerator. As seen, the price response is more negative the lower the implied supply elasticity is. On average, the bust price drop is about 10 percent larger when the financial accelerator is accounted for, while for the most restricted area this number is as high as 35 percent. While it would require a dynamic model to investigate the quantitative importance of the price-to-price feedback loop for the bust price drop, the results in the previous section clearly suggest that this would contribute to make the bust price drop in inelastic markets even greater relative to the elastic markets.

Pavlov and Wachter (2006) analyze the previous bust in US house prices in the 1990s. Both theoretically and empirically, they show that regions that were more exposed to aggressive lending instruments during the boom also experienced a larger price drop during the bust.[25] Based on the results of this paper, this can be attributed to a larger financial accelerator effect in more supply inelastic areas during the house price boom, possibly in combination with a price-to-price feedback loop.

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[25]Supporting results are found in a more recent article by the same authors, see Pavlov and Wachter (2011).

Figure 6: Bust price response for different supply elasticities
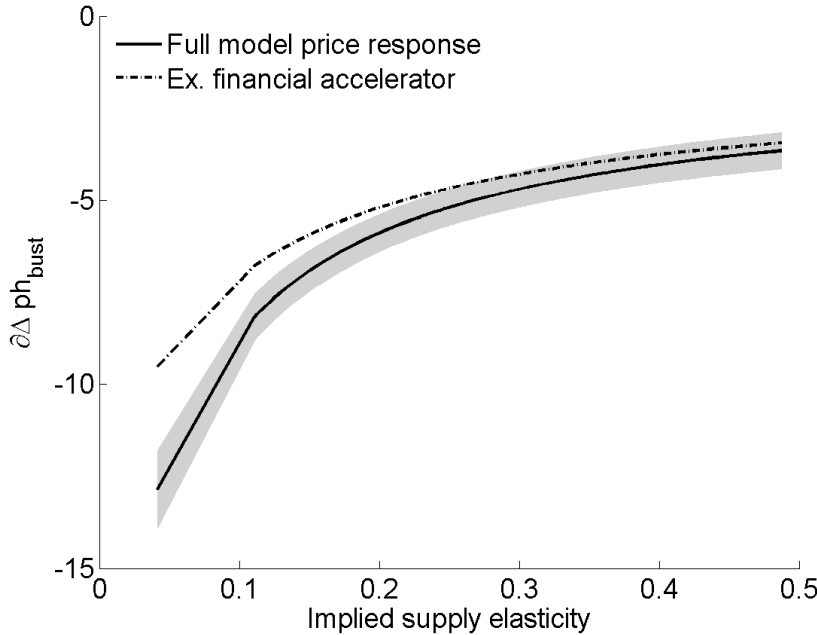
Note: This figure shows the bust period price response to a 1% shock to subprime lending per capita both with and without the financial accelerator in the model. It also shows the contribution coming from the boom period supply overhang. The confidence bounds are calculated using the delta method, see Appendices D and E.

5   Robustness

We have investigated how our results are affected by controlling for a range of different initial conditions. The set of controls includes log income per capita, the unemployment rate, log population, population density, poverty rates and a dummy for whether the MSA is situated in a state where lending is recourse. All variables, except the poverty rates for which we only have data from 1997, are measures as of 1996, i.e. prior to the start of the boom. Detailed results are reported in Table F.1 in Appendix F.

Adding these variables does not materially affect any of the coefficients in our model. Interestingly, the recourse dummy only enters significantly in the subprime equation, with a negatively signed and highly significant coefficient. This suggests that subprime lending was higher in areas where lending was non-recourse, which is consistent with recent evidence (Nam and Oh, 2013). The prevalence of non-recourse lending induces a put-option on the borrowers' side, since they can walk away from the house if they are "underwater" (Pavlov and Wachter, 2004, 2006). Nam and Oh (2013) study how the emergence of the originate-to-distribute model enabled banks to pass along risk to other investors, which reduced the incentive for screening and thereby contributed to a disproportionate increase in poor quality loans in non-recourse states and amplified the housing cycle in the 2000s. This is exactly the same as would be conjectured from our findings.

As one might be worried that the inference may be biased when normality is violated, we have also re-estimated our system excluding some large outliers. We excluded the

12 largest outliers, which is enough to ensure that normality of the system is satisfied (p-value = 0.1300). Excluding these areas has no major implications for our results. All coefficients are of a similar magnitude and are highly significant (see Table F.2 in Appendix F for details).

6 Conclusion

In this paper, we have analyzed the importance of supply restrictions and subprime lending for regional housing market developments through the recent US boom-bust cycle. Emphasis has been given to how housing markets with different supply elasticities respond to an increase in subprime lending. In particular, we have analyzed how price and quantity responses depend on the interaction between housing supply elasticities, a financial accelerator mechanism and a price-to-price feedback loop. The aim of the analysis has been to answer the following questions: How do restrictions on housing supply affect the housing market dynamics over a boom-bust cycle? And, is there evidence of financial accelerator and price-to-price expectation mechanisms? If so, how do these depend on supply restrictions and could they offer an explanation to the diverse results found in the earlier empirical literature?

Theoretically, we demonstrated that in a model without a financial accelerator, more restricted areas are predicted to see relatively large adjustments in prices following a positive demand shock, whereas areas with few restrictions on supply are expected to see large quantity adjustments. Both these forces should have a negative impact on house prices during the bust period. A baseline model without a price-credit interaction even suggests that the bust price response is independent of the supply elasticity. These theoretical conjectures are changed when we consider a model with a financial accelerator effect. First, restricted areas are expected to see an even larger price adjustment following a positive demand shock, since there is a relatively greater increase in collateral in these areas. Second, the difference in the quantity response across areas is expected to narrow, since the larger price acceleration in supply-inelastic markets has an additional stimulating effect on construction activity. Third, supply restricted areas are expected to be hit harder during the bust period.

To study these mechanisms empirically, we considered an econometric model including price, quantity and subprime credit equations. We found that house prices and credit are mutually reinforcing; tighter supply restrictions lead to a stronger financial accelerator, with additional positive effects on both prices and quantity. Although more supply restricted areas experience a relatively low quantity response for a given price increase, the stronger endogenous price acceleration in these areas partly dilutes the relation between supply restrictions and the total quantity response. In particular, we cannot reject an equal quantity response across all areas.

Further, in addition to the financial accelerator, we allowed lagged prices, interpreted as capturing adaptive expectations, to have an effect on prices, quantity and subprime lending. Our results suggest that adaptive price expectations primarily affect house prices through the lending channel, and not directly through housing demand or supply. Thus, an expectation of higher house prices is primarily materialized into increased demand for housing because banks loosen their lending standards and/or households increase their

credit demand.

In combination, these results suggest that one reason why more supply-restricted areas witnessed a greater price drop during the recent bust period is that they experienced a substantially larger credit boom, as a result of both a financial accelerator effect and a price-to-price expectation mechanism. Hence, these areas had a larger price overhang at the peak of the boom, while the quantity overhang was close to that of the less restricted areas.

We find that regulatory supply restrictions are more important than geographical supply restrictions. Hence, from a policy point of view, our results suggest that, in order to minimize the amplitude of a house price cycle, one should refrain from aggressive regulation of the housing supply. At least, if the amplitude of boom-bust cycles is a concern, a tighter regulatory environment for the construction sector should be accompanied by stricter credit market regulations.

In light of our results, a promising avenue for future research is to study these region-specific price acceleration mechanisms, while accounting for possible endogenous political changes in the regulatory framework through the boom-bust cycle. When more data become available, it will be particularly interesting to either consider the effect of changes in regulation in a dynamic panel model, or estimate separate time series models for individual MSAs.

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A Proofs for theory model

Proof of Proposition 2.1

In the model *without* a financial accelerator, the effects on house prices and housing supply of an increase in a demand shifter $\hat{v}_{0,i,t} \in v_{0,i,t}$ are given by:

$$\frac{\partial ph_{i,t}}{\partial \hat{v}_{0,i,t}}\bigg|^{M=B} = \frac{1}{1 - v_1 \varphi_i} > 0$$

$$\frac{\partial h_{i,t}}{\partial \hat{v}_{0,i,t}}\bigg|^{M=B} = \frac{\varphi_i}{1 - v_1 \varphi_i} > 0$$

As a result, the differences in the price and quantity responses for a *low* elasticity $(\underline{\varphi})$ and a *high* elasticity $(\overline{\varphi})$ market with $0 < \underline{\varphi} < \overline{\varphi} < \infty$ are given by:

1.

$$\frac{\partial ph_{i,t}}{\partial \hat{v}_{0,i,t}}\bigg|^{M=B}_{\varphi_i = \underline{\varphi}} - \frac{\partial ph_{i,t}}{\partial \hat{v}_{0,i,t}}\bigg|^{M=B}_{\varphi_i = \overline{\varphi}} = \frac{1}{1 - v_1 \underline{\varphi}} - \frac{1}{1 - v_1 \overline{\varphi}} = \frac{v_1 \left(\underline{\varphi} - \overline{\varphi} \right)}{\left(1 - v_1 \underline{\varphi} \right)\left(1 - v_1 \overline{\varphi} \right)} > 0$$

2.

$$\frac{\partial h_{i,t}}{\partial \hat{v}_{0,i,t}}\bigg|^{M=B}_{\varphi_i = \underline{\varphi}} - \frac{\partial h_{i,t}}{\partial \hat{v}_{0,i,t}}\bigg|^{M=B}_{\varphi_i = \overline{\varphi}} = \frac{\underline{\varphi}}{1 - v_1 \underline{\varphi}} - \frac{\overline{\varphi}}{1 - v_1 \overline{\varphi}} = \frac{\left(\underline{\varphi} - \overline{\varphi} \right)}{\left(1 - v_1 \underline{\varphi} \right)\left(1 - v_1 \overline{\varphi} \right)} < 0$$

To shift the demand curve back to its initial position in period $t+1$, we need a negative shock of size $\frac{\partial ph_{i,t}}{\partial \hat{v}_{0,i,t}}\big|^{M=B}$. But, even though I_t drops to zero when $ph_{i,t+1} \leq c_{0,i}$, it is evident from (3) and (4) that ph_{t+1} will be affected also through the effect that $\hat{v}_{0,i,t}$ has on h_t, i.e. the quantity overhang. The total effect on ph_{t+1} of reversing the shock is given by:

$$\frac{\partial ph_{i,t+1}}{\partial \hat{v}_{0,i,t}}\bigg|^{M=B} = - \frac{\partial ph_{i,t}}{\partial \hat{v}_{0,i,t}}\bigg|^{M=B} + v_1 \frac{\partial h_{i,t}}{\partial \hat{v}_{0,i,t}}\bigg|^{M=B} = - \frac{1}{1 - v_1 \varphi_i} + v_1 \frac{\varphi_i}{1 - v_1 \varphi_i} = -1$$

Proof of Proposition 2.2

Ruling out an explosive solution $(\eta \kappa_1 < 1 - v_1 \varphi_i)$, the effects on house prices, housing supply and borrowing of an increase in a demand shifter $\hat{v}_{0,i,t} \in \tilde{v}_{0,i,t} \subset v_{0,i,t}$ in the model *with* a financial accelerator are given by:

$$\frac{\partial ph_{i,t}}{\partial \hat{v}_{0,i,t}}\bigg|^{M=E} = \frac{1}{1 - v_1 \varphi_i - \eta \kappa_1} > \frac{1}{1 - v_1 \varphi_i} > 0$$

$$\frac{\partial h_{i,t}}{\partial \hat{v}_{0,i,t}}\bigg|^{M=E} = \frac{\varphi_i}{1 - v_1 \varphi_i - \eta \kappa_1} > \frac{\varphi_i}{1 - v_1 \varphi_i} > 0$$

$$\frac{\partial b_{i,t}}{\partial \hat{v}_{0,i,t}}\bigg|^{M=E} = \frac{\kappa_1}{1 - v_1 \varphi_i - \eta \kappa_1} > 0$$

Thus, introducing the financial accelerator, both the price and the quantity responses are greater than in the model without such effects. The differences in the price and the quantity responses for a *low* elasticity $(\underline{\varphi})$ and a *high* elasticity $(\overline{\varphi})$ market with $0 < \underline{\varphi} < \overline{\varphi} < \infty$ are given by:

1.

$$\left.\frac{\partial ph_{i,t}}{\partial \hat{v}_{0,i,t}}\right|_{\varphi_i=\underline{\varphi}}^{M=E} - \left.\frac{\partial ph_{i,t}}{\partial \hat{v}_{0,i,t}}\right|_{\overline{\varphi_i=\varphi}}^{M=E} = \frac{1}{1-v_1\underline{\varphi}-\eta\kappa_1} - \frac{1}{1-v_1\overline{\varphi}-\eta\kappa_1}$$

$$= \frac{v_1\left(\underline{\varphi}-\overline{\varphi}\right)}{\left(1-v_1\underline{\varphi}-\eta\kappa_1\right)\left(1-v_1\overline{\varphi}-\eta\kappa_1\right)} > 0$$

2.

$$\left.\frac{\partial h_{i,t}}{\partial \hat{v}_{0,i,t}}\right|_{\varphi_i=\underline{\varphi}}^{M=E} - \left.\frac{\partial h_{i,t}}{\partial \hat{v}_{0,i,t}}\right|_{\varphi_i=\overline{\varphi}}^{M=E} = \frac{\underline{\varphi}}{1-v_1\underline{\varphi}-\eta\kappa_1} - \frac{\overline{\varphi}}{1-v_1\overline{\varphi}-\eta\kappa_1}$$

$$= \frac{\left(\underline{\varphi}-\overline{\varphi}\right)\left(1-\eta\kappa_1\right)}{\left(1-v_1\underline{\varphi}-\eta\kappa_1\right)\left(1-v_1\overline{\varphi}-\eta\kappa_1\right)}$$

(a)

$$\left.\frac{\partial h_{i,t}}{\partial \hat{v}_{0,i,t}}\right|_{\varphi_i=\underline{\varphi}}^{M=E} - \left.\frac{\partial h_{i,t}}{\partial \hat{v}_{0,i,t}}\right|_{\varphi_i=\overline{\varphi}}^{M=E} < 0 \text{ iff } \eta\kappa_1 < 1$$

(b)

$$\left.\frac{\partial h_{i,t}}{\partial \hat{v}_{0,i,t}}\right|_{\varphi_i=\underline{\varphi}}^{M=E} - \left.\frac{\partial h_{i,t}}{\partial \hat{v}_{0,i,t}}\right|_{\varphi_i=\overline{\varphi}}^{M=E} = 0 \text{ iff } \eta\kappa_1 = 1$$

(c)

$$\left.\frac{\partial h_{i,t}}{\partial \hat{v}_{0,i,t}}\right|_{\varphi_i=\underline{\varphi}}^{M=E} - \left.\frac{\partial h_{i,t}}{\partial \hat{v}_{0,i,t}}\right|_{\varphi_i=\overline{\varphi}}^{M=E} > 0 \text{ iff } \eta\kappa_1 > 1$$

3.

$$\left.\frac{\partial b_{i,t}}{\partial \hat{v}_{0,i,t}}\right|_{\varphi_i=\underline{\varphi}}^{M=E} - \left.\frac{\partial b_{i,t}}{\partial \hat{v}_{0,i,t}}\right|_{\varphi_i=\overline{\varphi}}^{M=E} = \frac{\kappa_1}{1-v_1\underline{\varphi}-\eta\kappa_1} - \frac{\kappa_1}{1-v_1\overline{\varphi}-\eta\kappa_1}$$

$$= \frac{\kappa_1 v_1\left(\underline{\varphi}-\overline{\varphi}\right)}{\left(1-v_1\underline{\varphi}-\eta\kappa_1\right)\left(1-v_1\overline{\varphi}-\eta\kappa_1\right)} > 0$$

Following the same argument as in the *Proof of proposition 2.1*, the total fall in prices during the bust is given as:

$$\left.\frac{\partial ph_{i,t+1}}{\partial \hat{v}_{0,i,t}}\right|^{M=E} = -\left.\frac{\partial ph_{i,t}}{\partial \hat{v}_{0,i,t}}\right|^{M=E} + v_1\left.\frac{\partial h_{i,t}}{\partial \hat{v}_{0,i,t}}\right|^{M=E} = -\frac{1}{1-v_1\varphi_i-\eta\kappa_1} + v_1\frac{\varphi_i}{1-v_1\varphi_i-\eta\kappa_1}$$

$$= -\frac{1-v_1\varphi_i}{1-v_1\varphi_i-\eta\kappa_1} < 0$$

As a result, the difference in the bust price response for a *low* elasticity ($\underline{\varphi}$) and a *high* elasticity ($\overline{\varphi}$) market with $0 < \underline{\varphi} < \overline{\varphi} < \infty$ is given by:

$$
\frac{\partial ph_{i,t+1}}{\partial \hat{v}_{0,i,t}}\bigg|^{M=E}_{\varphi_i=\underline{\varphi}} - \frac{\partial ph_{i,t+1}}{\partial \hat{v}_{0,i,t}}\bigg|^{M=E}_{\varphi_i=\overline{\varphi}} = -\frac{1 - v_1\underline{\varphi}}{1 - v_1\underline{\varphi} - \eta\kappa_1} + \frac{1 - v_1\overline{\varphi}}{1 - v_1\overline{\varphi} - \eta\kappa_1}
$$

$$
= \frac{v_1\eta\kappa_1\left(\overline{\varphi} - \underline{\varphi}\right)}{\left(1 - v_1\underline{\varphi} - \eta\kappa_1\right)\left(1 - v_1\overline{\varphi} - \eta\kappa_1\right)} < 0
$$

B HMDA data calculations

As a part of the supervisory system, the US Congress mandated in 1975, through the Home Mortgage Disclosure Act (HMDA), that most banks in metropolitan areas disclose information on certain characteristics of the loan applications they have received during a calendar year. In 1989, the coverage was extended to include information on race, ethnicity, loan decisions, et cetera, at the applicant level. These data are available for the period 1990-2010, and we were able to collect data at the loan applicant level for the period 1996-2010, covering the recent US housing boom-bust cycle. The HMDA data have a wide coverage and are likely to be representative of lending in the US. For an excellent summary of the opportunities and limitations of the data, see the discussion in Avery et al. (2007). In 2010, the HMDA data covered 7923 home lending institutions and 12.95 million applications (see Avery et al. (2010)). In contrast, in the years prior to the housing collapse (the 2000-2006 period), the average number of applications reported in the registry was nearly 32 million.

While the data are available at the applicant level, the focus of our study is regional differences in US house price dynamics, and in particular the role of credit conditions in the recent boom-bust cycle. The individual data do have regional identifiers, which we have utilized to construct our data set. That said, due to definitional changes by the Census Bureau in the geographical composition of the different MSAs in 1993, 1999 and 2004, the data construction process was considerably complicated. To keep the geographical area spanned by the different MSAs constant and to remain consistent with the MSA definitions used in the Moody's data, we have relied on the 2004 definitions.

We limit ourselves to one-to-four family housing units, and follow the suggestion of Avery et al. (2007) and leave out small business loans from the calculations. That is, we drop all loans where information on the sex and race of both the applicant and the co-applicant is missing. We also noted some extremely large loan and income observations in the data that lead to unreasonable average income amounts as well as loan amounts. We suspect this is caused by reporting errors, and use the error list sent by HMDA to the reporting institutions to eliminate these from our sample. Information on the list for validity and syntactical edits is provided here: http://www.ffiec.gov/hmda/edits.htm. Detailed information on the error check list and how we implemented this is available upon request. Very few loans are in fact deleted from the data, but the average loan size as well as income figures are much more reasonable after this has been done.

Before 2004, the HMDA data contained no information on the lien status of the loan, which is important to avoid "double counting". To take account of this, we have followed an approach similar to Calhoun (2006). The approach may be described in two steps. In the first step, we follow Avery et al. (2007) and sort all observations in a given MSA and within a given year by certain person identifiers and a bank identifer (the respondent ID). The person identifiers include income of applicant, tract code, race of applicant, race of co-applicant, sex of applicant, sex of co-applicant and information on whether the property that the loan is secured against is an owner-occupied unit or not. If we get a match, we identify this as the same borrower and the smaller of the two loans is classified as the second lien (the "piggyback") and the larger is the first lien loan. We then exclude these observations from our selection sample. In the second step, we follow Calhoun (2006) and LaCour-Little et al. (2011) and perform a similar sorting and matching procedure,

only now we leave out the bank identifer. These observations are then removed from the sample, and we have three data sets: One with multi-loans as identified at step one, one with multi-loans as identified at step two and one containing only single loans. Finally, we match all these data sets and perform our calculations to generate variables at an MSA level. We deviate from previous papers in that we do not allow loans without income information to be included in a loan portfolio. The argument is that missing income information does not allow us to uniquely (to the extent it is possible without a social security number) identify the borrower. For the years 2004-2010, where we also have information on the lien status of the loan, we have performed a robustness check of the second liens as classified by our procedure, and we find a very high match. This is important to get a more precise measure of average LTI ratios and the number of loans originated in general. Finally, after correcting the data, we identify a loan as being a subprime loan if the bank extending the loan appears on the HUD Subprime and Manufactured Home Lender List: http://www.huduser.org/portal/datasets/manu.html.

C Reduced form representations

The baseline model

The reduced form representation of the baseline boom system (equations (7) and (8)) is given by:

$$\Delta ph_i^{Boom} = \frac{1}{A^B} \left[(\alpha_1 + \beta_{1,\Delta h}\alpha_2) + \boldsymbol{\beta}'_{1,\boldsymbol{x}}\boldsymbol{x}_i + \beta_{1,\Delta h}\boldsymbol{\beta}'_{2,\boldsymbol{z}}\boldsymbol{z}_i + \right] + u_{1,i}^B \tag{C.1}$$

$$\Delta h_i^{Boom} = \frac{1}{A^B} \left[(\alpha_1(\beta_{2,\Delta ph} + \beta_{2,\Delta ph \times \boldsymbol{Reg}}\boldsymbol{Reg}_i) + \alpha_2) \right.$$
$$\left. + (\beta_{2,\Delta ph} + \beta_{2,\Delta p \times \boldsymbol{Reg}}\boldsymbol{Reg}_i)\boldsymbol{\beta}'_{1,\boldsymbol{x}}\boldsymbol{x}_i + \boldsymbol{\beta}'_{2,\boldsymbol{z}}\boldsymbol{z}_i \right] + u_{2,i}^B \tag{C.2}$$

where $A^B = 1 - \beta_{1,\Delta h}(\beta_{2,\Delta p} + \beta_{2,\Delta h \times \boldsymbol{Reg}}\boldsymbol{Reg}_i)$, and:

$$u_{1,i}^B = \frac{1}{A^B}\left(\varepsilon_{\Delta ph,i} + \beta_{1,\Delta h}\varepsilon_{\Delta h,i}\right)$$

$$u_{2,i}^B = \frac{1}{A^B}\left((\beta_{2,\Delta ph} + \beta_{2,\Delta ph \times \boldsymbol{Reg}}\boldsymbol{Reg}_i)\varepsilon_{\Delta ph,i} + \varepsilon_{\Delta h,i}\right)$$

The extended model

The reduced form representation of the boom system with the subprime measure treated as endogenous (equations (7), (8), and (9)) is given by:

$$\Delta ph_i^{Boom} = \frac{1}{A^E}\left[(\alpha_1 + \beta_{1,\Delta h}\alpha_2 + \beta_{1,\Delta sp}\alpha_3) + \boldsymbol{\beta}'_{1,\boldsymbol{x}}\boldsymbol{x}_i + \beta_{1,\Delta h}\boldsymbol{\beta}'_{2,\boldsymbol{z}}\boldsymbol{z}_i + \beta_{1,\Delta sp}\boldsymbol{\beta}'_{3,\boldsymbol{w}}\boldsymbol{w}_i\right] + u_{1,i}^E \tag{C.3}$$

$$\Delta h_i^{Boom} = \frac{1}{A^E}\left[(\alpha_1(\beta_{2,\Delta ph} + \beta_{2,\Delta ph \times \boldsymbol{Reg}}\boldsymbol{Reg}_i) + \alpha_2(1 - \beta_{1,\Delta sp}\beta_{3,\Delta ph})\right.$$
$$+ \alpha_3\beta_{1,\Delta sp}(\beta_{2,\Delta ph} + \beta_{2,\Delta ph \times \boldsymbol{Reg}}\boldsymbol{Reg}_i))$$
$$+ (\beta_{2,\Delta ph} + \beta_{2,\Delta ph \times \boldsymbol{Reg}}\boldsymbol{Reg}_i)\boldsymbol{\beta}'_{1,\boldsymbol{x}}\boldsymbol{x}_i + (1 - \beta_{1,\Delta sp}\beta_{3,\Delta ph})\boldsymbol{\beta}'_{2,\boldsymbol{z}}\boldsymbol{z}_i$$
$$\left. + \beta_{1,sp}(\beta_{2,\Delta ph} + \beta_{2,\Delta ph \times \boldsymbol{Reg}}\boldsymbol{Reg}_i)\boldsymbol{\beta}'_{3,\boldsymbol{w}}\boldsymbol{w}_i\right] + u_{2,i}^E \tag{C.4}$$

$$\Delta sp_i^{Boom} = \frac{1}{A^E}\left[\beta_{3,\Delta ph}\alpha_1 + \beta_{1,\Delta h}\beta_{3,\Delta ph}\alpha_2 + (1 - \beta_{1,\Delta h}(\beta_{2,\Delta ph} + \beta_{2,\Delta ph \times \boldsymbol{Reg}}\boldsymbol{Reg}_i))\alpha_3\right.$$
$$\left. + \beta_{3,\Delta ph}\boldsymbol{\beta}'_{1,\boldsymbol{x}}\boldsymbol{x}_i + \beta_{1,\Delta h}\beta_{3,\Delta ph}\boldsymbol{\beta}'_{2,\boldsymbol{z}}\boldsymbol{z}_i + (1 - \beta_{1,\Delta h}(\beta_{2,\Delta ph} + \beta_{2,\Delta ph \times \boldsymbol{Reg}}\boldsymbol{Reg}_i))\boldsymbol{\beta}'_{3,\boldsymbol{w}}\boldsymbol{w}_i\right]$$
$$+ u_{3,i}^E \tag{C.5}$$

where $\boldsymbol{w} = \Delta y^{Bust}$, $A^E = 1 - \beta_{1,\Delta h}(\beta_{2,\Delta p} + \beta_{2,\Delta h \times \boldsymbol{Reg}}\boldsymbol{Reg}_i) - \beta_{3,\Delta p}\beta_{1,\Delta sp}$, and:

$$u_{1,i}^E = \frac{1}{A^E}\left(\varepsilon_{\Delta ph,i} + \beta_{1,\Delta h}\varepsilon_{\Delta h,i} + \beta_{1,\Delta sp}\varepsilon_{\Delta sp,i}\right)$$

$$u_{2,i}^E = \frac{1}{A^E}\left((\beta_{2,\Delta ph} + \beta_{2,\Delta ph \times \boldsymbol{Reg}}\boldsymbol{Reg}_i)\varepsilon_{\Delta ph,i} + (1 - \beta_{1,\Delta sp}\beta_{3,\Delta ph})\varepsilon_{\Delta h,i}\right.$$
$$\left. + \beta_{1,\Delta sp}(\beta_{2,\Delta ph} + \beta_{1,\Delta sp}\beta_{2,\Delta ph \times \boldsymbol{Reg}}\boldsymbol{Reg}_i)\varepsilon_{\Delta sp,i}\right)$$

$$u_{3,i}^E = \frac{1}{A^E}\left(\beta_{3,\Delta ph}\varepsilon_{\Delta ph,i} + \beta_{1,\Delta h}\beta_{3,\Delta ph}\varepsilon_{\Delta h,i} + (1 - \beta_{1,\Delta h}(\beta_{2,\Delta ph} + \beta_{2,\Delta ph \times \boldsymbol{Reg}}\boldsymbol{Reg}_i))\varepsilon_{\Delta sp,i}\right)$$

The bust equation may therefore be expressed in terms of the structural parameters of the boom system (equations (7) and (8)) by replacing (C.3) and (C.4) in ((10)):

$$\Delta ph_i^{Bust} = \mu$$

$$+ \gamma_{\Delta ph} \left[\frac{1}{A^E} \left[(\alpha_1 + \beta_{1,\Delta h}\alpha_2 + \beta_{1,\Delta sp}\alpha_3) + \boldsymbol{\beta}'_{1,\boldsymbol{x}}\boldsymbol{x}_i + \beta_{1,\Delta h}\boldsymbol{\beta}'_{2,\boldsymbol{z}}\boldsymbol{z}_i + \beta_{1,\Delta sp}\boldsymbol{\beta}'_{3,\boldsymbol{w}}\boldsymbol{w}_i \right] \right]$$

$$+ \gamma_{\Delta h} \left[\frac{1}{A^E} \left[(\alpha_1(\beta_{2,\Delta ph} + \beta_{2,\Delta ph\times\boldsymbol{Reg}}\boldsymbol{Reg}_i) + \alpha_2(1 - \beta_{1,\Delta sp}\beta_{3,\Delta ph}) \right. \right.$$

$$+ \;\; \alpha_3\beta_{1,\Delta sp}(\beta_{2,\Delta ph} + \beta_{2,\Delta ph\times\boldsymbol{Reg}}\boldsymbol{Reg}_i)) + (\beta_{2,\Delta ph} + \beta_{2,\Delta ph\times\boldsymbol{Reg}}\boldsymbol{Reg}_i)\boldsymbol{\beta}'_{1,\boldsymbol{x}}\boldsymbol{x}_i$$

$$+ \; (1 - \beta_{1,\Delta sp}\beta_{3,\Delta ph})\boldsymbol{\beta}'_{2,\boldsymbol{z}}\boldsymbol{z}_i + \beta_{1,sp}(\beta_{2,\Delta ph} + \beta_{2,\Delta ph\times\boldsymbol{Reg}}\boldsymbol{Reg}_i)\boldsymbol{\beta}'_{3,\boldsymbol{w}}\boldsymbol{w}_i \Big] \Big] + u_{4,i}^E$$

$$\text{(C.6)}$$

with:

$$u_{4,i}^E = \frac{1}{A^E} \left(\gamma_{\Delta ph}u_{1,i}^E + \gamma_{\Delta h}u_{2,i}^E \right) + e_i$$

D The analytical expressions for the response functions

The baseline model

In the baseline model (see (7) and (8)), the subprime measure is part of the vector $\boldsymbol{x}_i$. If we let the subprime measure be denoted Δsp_i, and also let $\beta_{1,\Delta sp} \in \boldsymbol{\beta_x}$ be the coefficient on the subprime measure in the house price equation, while remembering that $A^B = 1 - \beta_{1,\Delta h}(\beta_{2,\Delta ph} + \boldsymbol{\beta'_{2\Delta ph \times Reg}}\boldsymbol{Reg}_i)$, it is straightforward to show that the effect on house prices and supply during the boom, as well as prices during the bust, of an increase in subprime lending is given as:

$$\frac{\partial \Delta ph^{Boom}}{\partial \Delta sp_i} = \frac{1}{A^B}\beta_{1,\Delta sp} \tag{D.1}$$

$$\frac{\partial \Delta h^{Boom}}{\partial \Delta sp_i} = \frac{1}{A^B}\beta_{1,\Delta sp}(\beta_{2,\Delta ph} + \boldsymbol{\beta'_{2,\Delta ph \times Reg}}\boldsymbol{Reg}_i) \tag{D.2}$$

$$\frac{\partial \Delta ph^{Bust}}{\partial \Delta sp_i} = \gamma_{\Delta ph}\frac{\partial \Delta ph^{Boom}}{\partial \Delta sp_i} + \gamma_{\Delta h}\frac{\partial \Delta h^{Boom}}{\partial \Delta sp_i}$$

$$= \frac{1}{A^B}\beta_{1,\Delta sp}\left(\gamma_{\Delta ph} + \gamma_{\Delta h}(\beta_{2,\Delta ph} + \boldsymbol{\beta'_{2\Delta ph \times Reg}}\boldsymbol{Reg}_i)\right) \tag{D.3}$$

It is clear that both house prices and supply will increase during the boom following a shock to subprime lending, and prices will fall during the bust.

The extended model

In the extended model, we showed in Appendix C that:

$$A^E = 1 - \beta_{1,\Delta h}(\beta_{2,\Delta ph} + \boldsymbol{\beta'_{2\Delta ph \times Reg}}\boldsymbol{Reg}_i) - \beta_{3,\Delta ph}\beta_{1,\Delta sp}$$

i.e. if – hypothetically – all coefficient estimates are equal in the baseline and the extended model, then $A^E < A^B$ as long as prices affect subprime lending ($\beta_{3,\Delta ph} > 0$). This is due to the financial accelerator effect (as captured by $\beta_{3,\Delta ph}\beta_{1,\Delta sp}$). Again, it is straightforward to show that the effect on house prices and quantity during the boom, as well as prices during the bust, of an increase in subprime lending (now interpreted as a shock to $\varepsilon_{\Delta sp,i}$ in equation (9)) are given as:

$$\frac{\partial \Delta ph^{Boom}}{\partial \varepsilon_{\Delta sp,i}} = \frac{1}{A^E}\beta_{1,\Delta sp} \tag{D.4}$$

$$\frac{\partial \Delta h^{Boom}}{\partial \varepsilon_{\Delta sp,i}} = \frac{1}{A^E}\beta_{1,\Delta sp}(\beta_{2,\Delta ph} + \boldsymbol{\beta'_{2,\Delta ph \times Reg}}\boldsymbol{Reg}_i) \tag{D.5}$$

$$\frac{\partial \Delta ph^{Bust}}{\partial \varepsilon_{\Delta sp,i}} = \gamma_{\Delta ph}\frac{\partial \Delta ph^{Boom}}{\partial \varepsilon_{\Delta sp,i}} + \gamma_{\Delta h}\frac{\partial \Delta h^{Boom}}{\partial \varepsilon_{\Delta sp,i}}$$

$$= \frac{1}{A^E}\beta_{1,\Delta sp}\left(\gamma_{\Delta ph} + \gamma_{\Delta h}(\beta_{2,\Delta ph} + \boldsymbol{\beta'_{2\Delta ph \times Reg}}\boldsymbol{Reg}_i)\right) \tag{D.6}$$

If $\beta_{3,\Delta ph} = 0$ (no effect on subprime lending from higher house prices), then $A^E = A^B$ and we are back at the baseline model.

E Calculation of standard errors using the delta method

In general, if $G(\theta)$ is a function of coefficients, then we know from the delta method that the variance of $G(\theta)$ is:

$$Var(G(\theta)) = G'(\theta)\Sigma_\theta (G'(\theta))^T \tag{E.1}$$

where Σ_θ is the covariance matrix of the parameters in θ. This expression will be used throughout this appendix to derive the analytical expressions for all the variances used to construct the confidence bounds for the response functions in Section 4.

The baseline model

The calculations here are based on the expressions for the first derivatives derived in Appendix D. The calculations are done to construct the confidence intervals used in the figures for the response functions in Section 4.1, where we condition on the implied supply elasticity.

Standard errors for boom price response

From (D.1), we have that:

$$G_i(\theta^{\Delta ph_i^{Boom}} \mid el_i) = \frac{\partial \Delta ph^{Boom}}{\partial \Delta sp_i} = \frac{\beta_{1,\Delta sp}}{A^B}$$

with $A^B = 1 - \beta_{1,\Delta h} el_i$ and $\theta^{\Delta ph^{Boom}} = (\beta_{1,\Delta sp}, \beta_{1,\Delta h})$. The implied elasticity, el_i, is given as $el_i = \beta_{2,\Delta ph} + \beta_{2,\Delta ph \times wrl} wrl_i + \beta_{2,\Delta ph \times una} una_i$. The vector of derivatives of $G_i(\theta^{\Delta ph_i^{Boom}} \mid el_i)$ is given as:

$$G_i'(\theta^{\Delta ph_i^{Boom}} \mid el_i) = \left(\frac{1}{A^B}, \frac{\beta_{1,\Delta sp} el_i}{A^{B2}} \right) \tag{E.2}$$

Using (E.1), we can then calculate the variance of $G_i(\theta^{\Delta ph_i^{Boom}} \mid el_i)$.

Standard errors for boom supply response

From (D.2), we have that:

$$G_i(\theta^{\Delta h_i^{Boom}} \mid el_i) = \frac{\partial \Delta h^{Boom}}{\partial \Delta sp_i} = \frac{\beta_{1,\Delta sp} el_i}{A^B}$$

with $A^B = 1 - \beta_{1,\Delta h} el_i$ and $\theta^{\Delta ph^{Boom}} = (\beta_{1,\Delta sp}, \beta_{1,\Delta h})$. We then find that the vector of derivatives of $G_i(\theta^{\Delta h_i^{Boom}} \mid el_i)$ is given as:

$$G_i'(\theta^{\Delta h_i^{Boom}} \mid el_i) = \left(\frac{el_i}{A^B}, \frac{\beta_{1,\Delta sp} el_i^2}{A^{B2}} \right) \tag{E.3}$$

We can again use the expression in (E.1) to calculate the variance of $G_i(\theta^{\Delta h_i^{Boom}} \mid el_i)$.

The extended model

The calculations below are based on expressions (D.4)–(D.6). The analytical expressions derived here are used to construct the confidence intervals used in the figures for the response functions in the extended model, see Section 4.2.

Standard errors for boom price response

From (D.4), we have that:

$$G_i(\theta^{\Delta ph_i^{Boom}} \mid el_i) = \frac{\partial \Delta ph_i^{Boom}}{\partial \varepsilon_{\Delta sp,i}} = \frac{\beta_{1,\Delta sp}}{A^E}$$

with $A^E = 1 - \beta_{1,\Delta h}el_i - \beta_{3,\Delta ph}\beta_{1,\Delta sp}$ and $\theta^{\Delta ph^{Boom}} = (\beta_{1,\Delta sp}, \beta_{1,\Delta h}, \beta_{3,\Delta ph})$. The vector of derivatives of $G_i(\theta^{\Delta ph_i^{Boom}} \mid el_i)$ is given as:

$$G_i'(\theta^{\Delta ph_i^{Boom}} \mid el_i) = \left(\frac{1 - \beta_{1,\Delta h}el_i}{A^{E2}}, \frac{\beta_{1,\Delta sp}el_i}{A^{E2}}, \frac{\beta_{1,\Delta sp}^2}{A^{E2}} \right) \tag{E.4}$$

Using the expression in (E.1), we can derive the variance of $G_i(\theta^{\Delta ph_i^{Boom}} \mid el_i)$.

Standard errors for boom supply response

From (D.5), we have that:

$$G_i(\theta^{\Delta h_i^{Boom}} \mid el_i) = \frac{\partial \Delta h_i^{Boom}}{\partial \varepsilon_{\Delta sp,i}} = \frac{\beta_{1,\Delta sp}el_i}{A^E}$$

with $A^E = 1 - \beta_{1,\Delta h}el_i - \beta_{3,\Delta ph}\beta_{1,\Delta sp}$ and $\theta^{\Delta ph^{Boom}} = (\beta_{1,\Delta sp}, \beta_{1,\Delta h}, \beta_{3,\Delta ph})$. We find that the vector of derivatives of $G_i(\theta^{\Delta h_i^{Boom}} \mid el_i)$ is given as:

$$G_i'(\theta^{\Delta h_i^{Boom}} \mid el_i) = \left(\frac{el_i(1 - \beta_{1,\Delta h}el_i)}{A^{E2}}, \frac{\beta_{1,\Delta sp}el_i^2}{A^{E2}}, \frac{\beta_{1,\Delta sp}^2 el_i}{A^{E2}} \right)$$

We then use the expression in (E.1) to calculate the variance of $G_i(\theta^{\Delta h_i^{Boom}} \mid el_i)$.

Standard errors for bust price response

The derivative of the bust price response conditional on the price and quantity responses in the boom is given as:

$$G_i'(\theta^{\Delta ph_i^{Bust}} \mid el_i) = \gamma_{\Delta ph} \, G_i'(\theta^{\Delta ph_i^{Boom}} \mid el_i) + \gamma_{\Delta h} \, G_i'(\theta^{\Delta h_i^{Boom}} \mid el_i)$$

where $\theta^{\Delta ph^{Bust}} = (\gamma_{\Delta ph}, \gamma_{\Delta h})$. The vector of derivatives for $G_i'(\theta^{\Delta ph_i^{Bust}} \mid el_i)$ is given as:

$$G_i'(\theta^{\Delta ph_i^{Bust}} \mid el_i) = \left(G_i'(\theta^{\Delta ph_i^{Boom}} \mid el_i) \,, G_i(\theta^{\Delta h_i^{Boom}} \mid el_i) \right) \tag{E.5}$$

Again, expression (E.1) is used to calculate the variance of $G(\theta^{\Delta h^{Boom}})$.

F Results from robustness

Table F.1: The boom period model including a financial accelerator and adaptive expectations including various controls, 2000–2006

Variable	Δph_{Boom}		Δh_{Boom}		Δsp_{Boom}	
	Coefficient	t-value	Coefficient	t-value	Coefficient	t-value
Constant	6.696	2.763	0.384	2.541	-12.074	8.989
Δh_{Boom}	-12.244	4.006	*	*	*	*
Δph_{Boom}	*	*	0.559	6.746	0.408	2.055
$una \times \Delta ph_{Boom}$	*	*	-0.192	2.670	*	*
$wrl \times \Delta ph_{Boom}$	*	*	-0.566	4.568	*	*
$\Delta ph_{Pre-Boom}$	-0.234	0.434	-0.228	0.835	1.377	2.929
$una \times \Delta ph_{Pre-Boom}$	*	*	0.247	0.868	*	*
$wrl \times \Delta ph_{Pre-Boom}$	*	*	0.622	1.070	*	*
Δsp_{Boom}	0.742	4.914	*	*	*	*
Δy_{Boom}	5.153	4.059	*	*	1.623	3.376
Δcc_{Boom}	*	*	-0.188	4.349	*	*
una	*	*	-0.142	2.044	*	*
wrl	*	*	0.023	0.200	*	*
Denial share$_{1996}$	*	*	*	*	0.524	1.219
LTI_{1996}	*	*	*	*	1.255	6.245
Controls:						
$lycap_{1996}$	0.021	0.051	-0.121	2.875	0.061	0.148
$lpop_{1996}$	-0.158	2.618	0.004	0.496	0.221	4.258
$pop.den_{1996}$	-0.000	0.151	-0.000	1.400	-0.000	1.258
u_{1996}	-2.598	0.780	-1.602	5.116	2.004	0.797
$Recourse_{1996}$	0.122	1.259	-0.013	0.888	-0.291	3.042
$Poverty_{1997}$	0.717	0.980	0.025	0.227	-0.617	0.930
Normality test	p-value= 0.0075					
Observations	242					

Note: The table reports the FIML estimates of the boom system (7)–(9). The following abbreviations apply: h is the log housing stock, ph is log house prices, sp is log cumulative subprime originations per capita, y is log household income, cc is log construction wages, una is the geographical restriction index of Saiz (2010), wrl is the regulatory index of Gyourko et al. (2008) adjusted for una and normalized to range between 0 and 1, LTI_{1996} is the loan-to-income ratio in 1996 and Denial share$_{1996}$ is the denial share in 1996. All variables are nominal, and all variables except the subprime variable are in percentage changes. Δ is a difference operator. An asterisk indicates that the variable is not included in the equation under consideration. The reported t-values are measured in absolute terms. The normality test is the Doornik and Hansen (2008) normality test for multivariate normality.

Table F.2: The boom period model including a financial accelerator and adaptive expectations excluding large outliers, 2000–2006

Variable	Δph_{Boom}		Δh_{Boom}		Δsp_{Boom}	
	Coefficient	t-value	Coefficient	t-value	Coefficient	t-value
Constant	3.899	6.808	-0.039	0.843	-10.096	24.976
Δh_{Boom}	-9.988	5.785	*	*	*	*
Δph_{Boom}	*	*	0.523	7.378	0.423	2.189
$una \times \Delta ph_{Boom}$	*	*	-0.151	2.247	*	*
$wrl \times \Delta ph_{Boom}$	*	*	-0.596	5.204	*	*
$\Delta ph_{Pre-Boom}$	-0.451	1.072	-0.080	0.317	1.851	3.774
$una \times \Delta ph_{Pre-Boom}$	*	*	-0.009	0.030	*	*
$wrl \times \Delta ph_{Pre-Boom}$	*	*	0.471	0.908	*	*
Δsp_{Boom}	0.572	7.565	*	*	*	*
Δy_{Boom}	3.928	5.356	*	*	1.708	3.612
Δcc_{Boom}	*	*	-0.189	4.483	*	*
una	*	*	-0.141	1.912	*	*
wrl	*	*	0.029	0.269	*	*
Denial share$_{1996}$	*	*	*	*	0.917	2.019
LTI$_{1996}$	*	*	*	*	1.851	3.774
Normality test	p-value= 0.1300					
Observations	230					

Note: The table reports the FIML estimates of the boom system (7)–(9). The following abbreviations apply: h is the log housing stock, ph is log house prices, sp is log cumulative subprime originations per capita, y is log household income, cc is log construction wages, una is the geographical restriction index of Saiz (2010), wrl is the regulatory index of Gyourko et al. (2008) adjusted for una and normalized to range between 0 and 1, LTI_{1996} is the loan-to-income ratio in 1996 and Denial share$_{1996}$ is the denial share in 1996. All variables are nominal, and all variables expect the subprime variable are in percentage changes. Δ is a difference operator. An asterisk indicates that the variable is not included in the equation under consideration. The reported t-values are measured in absolute terms. The normality test is the Doornik and Hansen (2008) normality test for multivariate normality.