WORKING PAPER

CHANGING SUPPLY ELASTICITIES AND REGIONAL HOUSING BOOMS*

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Abstract

Recent developments in US house prices mirror those of the 1996-2006 boom, but the recovery in construction activity has been weak. Using data for 254 US metropolitan areas, we show that housing supply elasticities have fallen markedly in recent years. Housing supply elasticities have declined more in areas where land-use regulation has tightened the most, and in areas that experienced the sharpest housing busts. A lowering of the housing supply elasticity implies a stronger price responsiveness to demand shocks, whereas quantity reacts less. Consistent with this, we find that an expansionary monetary policy shock has a considerably stronger effect on house prices during the recent recovery than during the previous housing boom. At the same time, building permits respond less.

Keywords: *House prices; Heterogeneity; Housing supply elasticities; Monetary policy*

JEL classification: *C23, E32, E52, R31*

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1 Introduction

At the end of 2017, nominal US house prices were almost ten percent above the pre-recession peak. Despite the strong rise in house prices, construction activity has remained low and is considerably weaker than during the previous housing boom. A similar pattern is evident at the regional level. We document that this is related to a recent decline in housing supply elasticities. Furthermore, we argue that there are large regional differences in the extent of the decline. Against this background, we ask the following questions: (i) How does the decline in housing supply elasticities impact house price volatility and the transmission of housing demand shocks?; and (ii) What factors have contributed to changing housing supply elasticities?

We consider a quarterly panel data set covering 254 US Metropolitan Statistical Areas (MSAs), spanning the previous boom episode (1996–2006) and the recent recovery (2012–2017). Our analysis is confined to the two boom periods. While housing busts are interesting to analyze, there are two main reasons why we focus on boom episodes. First, our main interest is to study the different dynamics across similar housing episodes. Second, the durability of housing entails that housing supply is rigid downwards (Glaeser and Gyourko 2005), implying that the elasticity should fall towards zero in a bust. Since this should hold in all markets, local-specific factors, such as differences in topography and housing market regulation, should not matter for the responsiveness of housing supply during a bust. For each of the sub-samples, we estimate MSA-specific housing supply elasticities, using building permits as the dependent variable. The housing supply elasticity is computed as the coefficient on house prices, controlling for numerous MSA-specific variables that may affect housing supply. This exercise is non-trivial for at least two reasons. First, there are large regional variations. Second, there is likely reverse causality between construction activity and house prices.

With respect to regional variations, theory suggests that local differences in topography and regulation should impact housing supply elasticities. We take this into account by interacting house prices with the index of topographical constraints calculated by Saiz (2010) and with the index of regulatory restrictions from Gyourko et al. (2008). To deal with reverse causality, we use an instrumental variable (IV) approach. Our identification problem requires separating housing demand from housing supply. We consider two instruments for house prices that we argue lead to shifts in housing demand, but that do not shift housing supply. The first instrument exploits variation in crime rates across MSAs and over time, compiled by the Federal Bureau of Investigation (FBI). Given the negative impact crime can have on society, crime can be viewed as a negative amenity (Pope and Pope 2012). Crime rates should therefore capture exogenous variations in (negative) amenities that drive house price changes both across and within MSAs over time. The second instrument is real personal disposable income. Income is one of the main determinants of housing and consumption demand in standard macro and housing models (Dougherty and Van Order 1982, Buckley and Ermisch 1983, Meen 1990, Muellbauer and Murphy 1997, Meen 2001, 2002, Duca et al. 2011), but typically does not affect housing supply directly. Thus, from a theoretical point of view, this instrument should satisfy both the relevance and exogeneity conditions.

Our IV-estimates suggest that housing supply elasticities have declined. A direct implication of lower supply elasticities is that a given change in demand should have a stronger effect on house prices. We explore the relevance of this conjecture through the use of exogenous monetary policy shocks. Following a recent strand of the literature, we use high-frequency data to identify unexpected changes in the Fed policy rate (see e.g, Gürkaynak et al. 2005, Gertler and Karadi 2015, Nakamura and Steinsson 2018). The high-frequency identified (HFI) shocks isolate news about future policy actions that are orthogonal to changes in economic and financial variables. We then use a local projection instrumental variable approach (Jordà et al. 2015, Ramey 2016, Stock and Watson 2018) to explore how monetary policy shocks affect house prices and permits in the two booms.

Our results show considerable heterogeneity in responses across local housing markets. We estimate a substantially greater response in house prices to a monetary policy shock in supply-inelastic markets than in areas with an elastic supply. This holds true for both boom periods. We also document a substantial increase in the responsiveness of house prices to monetary policy shocks in recent years. In particular, our results suggest that for a metro area with a median housing supply elasticity, an exogenous monetary policy shock that lowers the interest rate by one percentage point led to an increase in real house prices of about ten percent after four years during the 1996-2006 boom. For the 2012-2017 recovery, the estimated response is 16 percent. Consistent with this, we find that building permits today increase about three percentage points less in response to the monetary policy shock.

We also find that there are regional differences in how much elasticities have declined. There are several reasons why housing supply elasticities may differ across areas and change over time (Green et al. 2005), including changes in regulation, demographics, and in expectations about future demand and house prices. In a recent study, Herkenhoff et al. (2018) show that there have been substantial changes in residential land-use regulation in most US states over time. Using their measure of time-varying land-use regulation, we find that elasticities have declined the most in areas where regulation has tightened more. Our results also suggest a larger decline in elasticities in areas that experienced the largest decline in house prices at the end of the previous decade. We interpret this as evidence that the fear of a new bust has led developers to be less price-responsive than before.

The results in this paper relate to several strands of the literature. First, a vast number of papers has emphasized local differences in housing supply elasticities as a central driver of cross-sectional variation in US house price developments (see e.g., Green et al. 2005, Gyourko et al. 2008, Saiz 2010, Huang and Tang 2012, Glaeser et al. 2014, Anundsen and Heebøll 2016). This literature has used time-invariant measures of housing supply elasticities to explore cross-sectional variation over the course of a boom-bust cycle, finding that supply-inelastic areas experience stronger house price booms than areas with an elastic housing supply. Our results are consistent with this view, but go a step further by showing that housing supply elasticities may change over time even within the same local market. This contributes to affect local – and possibly aggregate – house price volatility over time.

Second, there is a growing literature looking at the nexus between monetary policy and house prices (see e.g., Iacoviello 2005, Del Negro and Otrok 2007, Jarocinski and Smets 2008, Jordà et al. 2015, Williams 2011, 2015). These papers focus on the aggregate effects on house prices, which masks potential heterogeneity across regional housing markets. One exception is Aastveit and Anundsen (2017), who study the asymmetric effects of monetary policy on regional house prices for a sample ending in 2007Q4. We add to this literature by documenting non-trivial heterogeneous responses of regional house prices to a common monetary policy shock for both the 1996-2006 boom and the 2012-2017 boom. Furthermore, we document a sizeable drop in housing supply elasticities over time, which makes house prices even more responsive to monetary policy shocks today. Paul (2019) finds that the transmission of monetary policy to financial variables, such as stock prices and house prices, has become stronger over time. Our work can provide an economic interpretation of these findings: due to the lowering of housing supply elasticities, an aggregate shock that raises housing demand is absorbed mostly by house prices rather than through an increase in quantity.

Herkenhoff et al. (2018) argue that the stronger tightening of residential land-use regulation in highly productive states, particularly California and New York, has restricted the available land for housing and commercial use, raised house prices, reduced capital and labor reallocation, resulting in a substantial decrease in output and productivity. In a similar vein, Ganong and Shoag (2017) find that the decline in income convergence and migration rates across states since the 1980s can – at least partly – be attributed to tight land-use regulation and rising house prices in high-income states. Hsieh and Moretti (2019) document that stringent housing restrictions in highly-productive areas, such as New York and San Francisco Bay Area, result in significant output costs in the form of spatial misallocation of labor across US cities. In addition, Glaeser and Gyourko (2018) posit that highly regulated areas are characterized by higher house prices and smaller population growth relative to the level of demand. Our results relate to this literature by documenting that the tightening of land-use regulation has resulted in a lower supply elasticity, which in turn amplifies the responsiveness of house prices to demand shocks.

Our results are robust along several dimensions. We show that the decline in housing supply elasticities is evident when: (i) employing a Bartik-type instrumental variable approach; (ii) using total crime rates (sum of property crime and violent crime) as the crime variable instrument; (iii) using permit intensity as the dependent variable to allow the dynamics in permits to differ according to the existing stock of houses; (iv) replacing the measures of topographical and regulatory constraints with a summary measure of supply restrictions to account for the possibility that these two indicators might be correlated; and (v) controlling for mortgage originations to assess the impact on the housing supply response of subdued credit developments since the Great Recession. Finally, our results are robust to estimating supply elasticities using 10-year and 15-year rolling windows.

The rest of the paper proceeds as follows. In the next section, we offer a descriptive analysis of the housing boom in the 2000's and the ongoing boom. In Section 3, we describe the data and some stylized facts about the US housing cycle over the past 20 years. We discuss our econometric approach and estimate local housing supply elasticities for the two boom periods in Section 4. In Section 5, we analyze how changing supply elasticities affects housing market dynamics. In Section 6, we explore the factors that have led to declining housing supply elasticities. Robustness checks and alternative explanations for the disconnect between house prices and housing supply are discussed in Section 7. Section 8 concludes the paper.

2 The 1996-2006 boom versus the 2012-2017 recovery

At the national level, real US house prices have increased by more than 26 percent since the beginning of the housing recovery in mid-2012. The dynamics of real house prices during the recovery is similar to that of the previous housing boom. This is illustrated in the upper left panel of Figure 1, where we plot real house prices for both the 1996-2006 boom (red line) and the 2012-2017 recovery (blue line). We have scaled the price index so that it takes a value of 100 at the beginning of each period. The horizontal axis shows quarters around the beginning of the two booms, while the vertical line at zero is the starting point of both booms. In the upper right panel, we perform the same exercise when deflating house prices by per capita income. Remarkably, the current boom looks far stronger relative to income than the previous boom.[1] Although our house price index is a weighted repeat-sales index, measuring average price changes in repeat sales or refinancings on the same properties, we observe the same pattern in house prices across booms for new homes (Figure D.2 in Appendix D).

Despite similar – or even stronger – developments in house prices, housing supply has grown substantially less during the current boom (lower panel of Figure 1). While the cumulative increases in total building permits and housing starts were roughly 60 percent over the first five to six years of the previous boom, the cumulative increase between 2012 and 2017 has been around 16 percent. This holds true for both single-family and multi-family units, although the multi-family segment has recovered somewhat faster (Figure D.3 in Appendix D). Our measure of housing supply is building permits. Nevertheless, similar developments have been seen for existing homes available for sale (Figure D.4).[2]

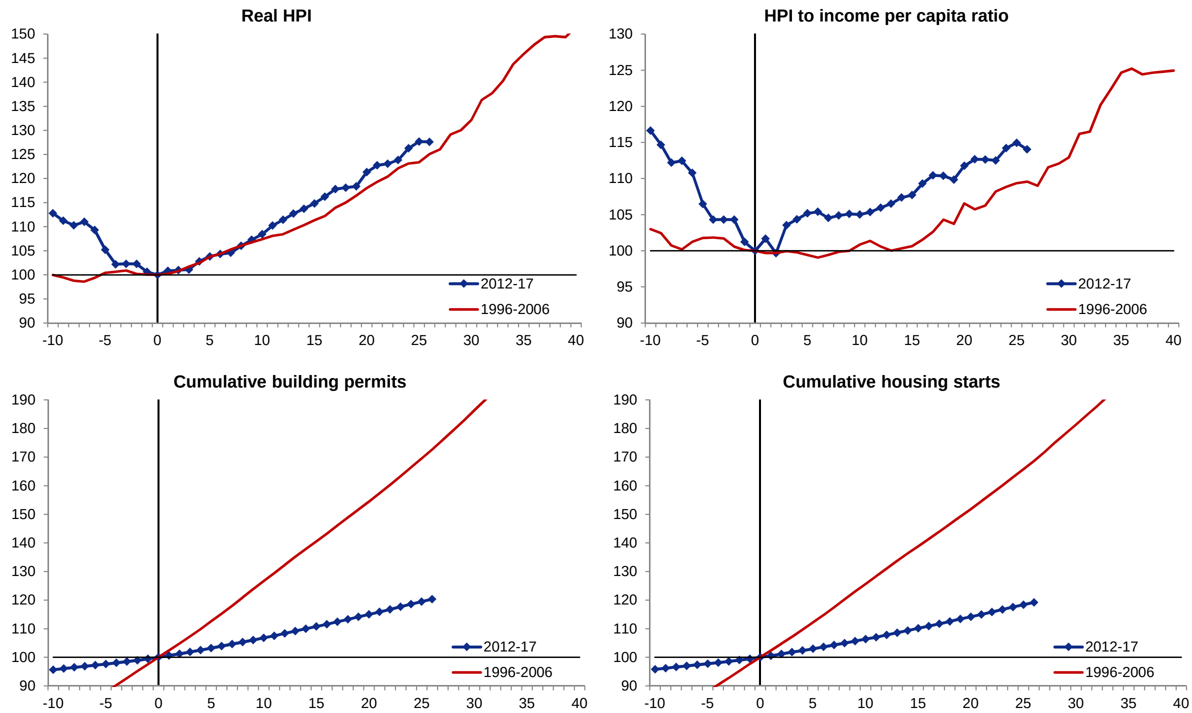
Housing is characterized by important regional heterogeneities (Ferreira and Gyourko 2012). We use MSA-level data and break the sample into quartiles of the cumulative house price change between 1996 and 2006. We define *Low HPI* MSAs as the areas belonging to the first quartile, while *High HPI* MSAs refers to the fourth quartile. We then compare the evolution of house prices relative to income and permits across the two booms (Figure 2). The red lines illustrate developments for the High HPI group, and green lines for the Low HPI group. To distinguish

[1]The strong developments in house prices relative to income per capita can be partially attributed to subdued income and consumption growth, as illustrated in Figure D.1 in Appendix D.

[2]In the current housing recovery, there has been a close link between new residential construction and the supply of existing homes listed for sale. With fewer new homes to choose from, many homeowners considering upgrading have chosen to remain in their current homes, and therefore have not listed them for sale. This has prevented other homeowners from upgrading as well, limiting the number of existing homes available for sale even further. Despite rising house prices in both the new and existing home segments, this 'vicious circle' between limited new homes in the market leading to a tight supply of existing homes for sale has been the norm in the current boom (Rappaport 2018).

between the two periods, we use dotted lines for the 1996-2006 period and solid lines for the 2012-2017 period. Mirroring the aggregate picture, house prices relative to income per capita have increased more during the current boom for both groups. At the same time, this ratio has increased most for the High HPI MSAs. In contrast, permits have progressed at a sluggish pace during the current recovery, with a slightly weaker expansion in High HPI MSAs.

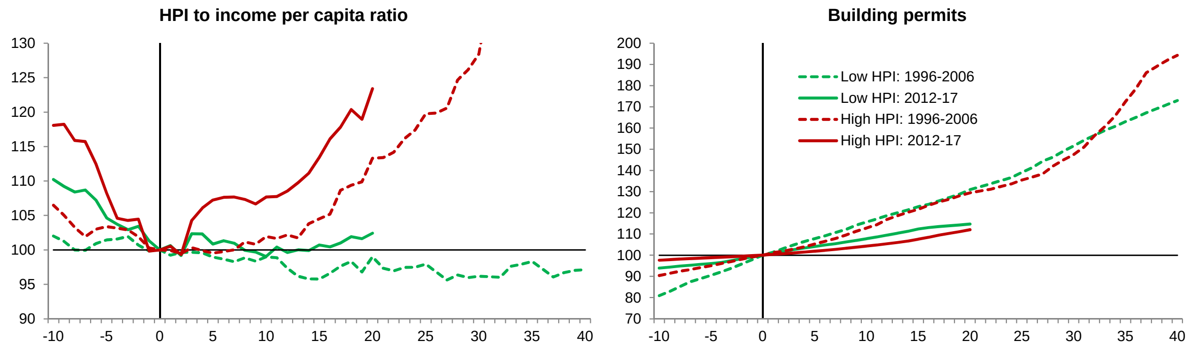
Figure 1: House price developments across booms



Sources: Bureau of Economic Analysis, Census Bureau, Federal Housing Finance Agency, and authors' calculations.
Notes: The figure shows developments in real house prices, house prices relative to income per capita, building permits, and housing starts during 1996q4–2006q4 (red solid line) and 2012q3–2017q4 (blue line with markers). The series are scaled such that they take a value of 100 at the beginning of both periods. The horizontal axis shows quarters around the beginning of the two booms, and the vertical line at zero is the starting point of both booms.

The marked differences in housing market developments across metropolitan areas highlight the importance of studying regional markets. The use of disaggregated data follows the most recent housing market literature, which tends to look at the housing market as a collection of several markets that differ not only geography but also by other attributes – see Piazzesi and Schneider (2016) for a survey.

Figure 2: Housing indicators for MSA groups across housing booms



Sources: Bureau of Economic Analysis, Census Bureau, Federal Housing Finance Agency, Moody's Analytics, and authors' calculations.
Notes: The figure shows developments in house prices relative to income per capita and permits for 1996q4–2006q4 (dotted lines) and 2012q3–2017q4 (solid lines). *Low HPI* MSAs (green) are the areas that recorded the smallest cumulative growth in house prices over 1996-2006, as measured by the first quartile, whereas *High HPI* MSAs (red) refers to the fourth quartile. The series are scaled such that they take a value of 100 at the beginning of each period. The horizontal axis shows quarters around the beginning of the two booms, and the vertical line at zero is the starting point of both booms.

3 Data and housing market cycles

3.1 Data

We use quarterly data for a panel of 254 MSAs between 1996 and 2017. The sample covers more than 80 percent of US income and population. Our MSA definitions follow the new delineations issued by the Office of Management and Budget (OMB), based on the 2010 Census. The MSA data on housing supply encompass building permits, housing starts, and the housing stock. In addition, we have data on house prices, and controls for macroeconomic, financial and socio-demographic conditions: personal disposable income, unemployment rates, mortgage originations, population, crime rates, dependency ratio (ratio of people younger than 15 or older than 64 relative to those aged 15-64), and the fraction of Blacks and Hispanics relative to the total population. We also use wages and salaries in the construction sector to proxy builders' costs. This series is available only at the state level. We deflate all nominal macroeconomic series with the MSA-level consumer price index (CPI). The MSA data have been provided by Moody's Analytics, with the original sources of the data coming mainly from the Census Bureau, Bureau of Economic Analysis (BEA), Bureau of Labor Statistics (BLS), and Federal Housing Finance Agency (FHFA). The exception is the crimes rates, which we compiled from publicly available reports from the FBI. A full list of variables, sources, and descriptive statistics is provided in Appendix B.

We control for regional differences in supply restrictions with two indices, which vary only at the cross-sectional level. First, we measure topographical supply restrictions with the UNAVAL index by Saiz (2010). UNAVAL measures MSA-level geographical land availability constraints. Saiz (2010) uses GIS and satellite information over 1970-2000 to calculate the share of land in a 50 kilometer radius of the MSA main city center 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. Saiz (2010) finds that metropolitan areas that are more inelastic are typically more land constrained. Second, we measure regulatory constraints with the Wharton Regulatory Land Use Index (WRLURI) from Gyourko et al. (2008). WRLURI measures the stringency of local zoning laws, i.e. the time and financial cost of acquiring building permits and constructing a new home. It is based on a nationwide survey in 2005, and on a separate study of state executive, legislative and judicial activity.[3]

3.2 Housing market cycles

To date booms and busts over the housing cycle, we analyze peaks and troughs in real house prices at the median.[4] For ease of illustration, we plot the national house price index, together with the median, the 10^{th} and 90^{th} percentiles of the house price distribution at the MSA level (Figure 3). We detect three phases of the housing cycle: a strong boom from 1996 until 2006, followed by a severe bust lasting until 2012.[5] By 2012, a new boom (the ongoing recovery) had started. With our data set, we cannot identify either a boom or a bust over 1986-1996. Instead, we observe significant heterogeneity across MSAs over this period; the MSAs at the bottom of the house price distribution recorded a steady increase in house prices, while the MSAs at the top saw the opposite dynamics. At the median real house prices remained relatively stable over that ten-year period.[6]

All of the MSAs experienced increasing house prices during the 1996-2006 boom, but dispersion was high; house prices increased by 17 percent, on average, for the MSAs belonging

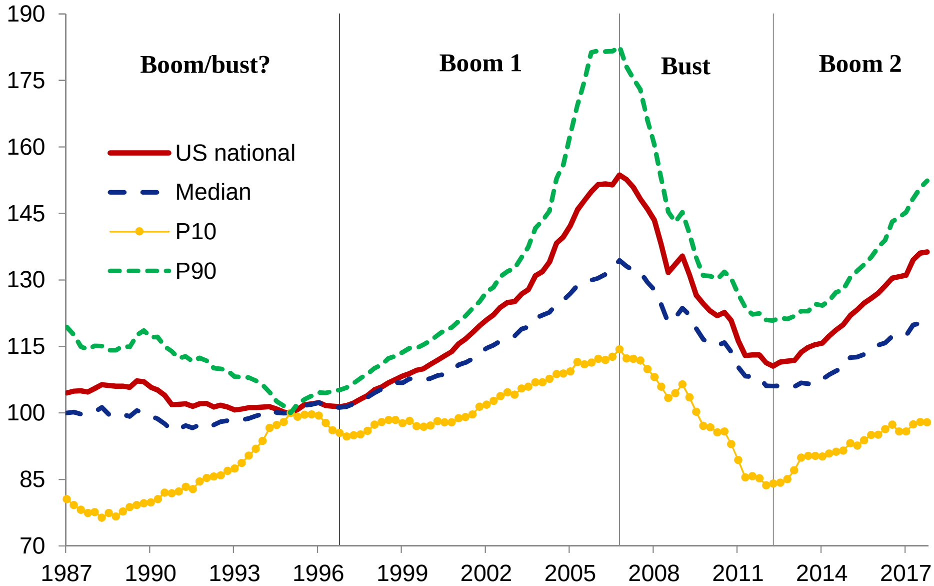
[3]This index is based on 11 sub-indices measuring different types of complications and regulations in the process of getting a building permit. WRLURI is available at a town (or city) level, which we have aggregated to the MSA level using the sample probability weights of Gyourko et al. (2008).

[4]Given that our sample of 254 MSAs includes some areas with large variations in prices, we look at the median, rather than the mean as in Glaeser et al. (2008). The median minimizes the effect that outliers have on dating the housing cycles. We track the evolution in the median real house price index over time, which does not mean necessarily that we track the same MSA over time. Alternative approaches to ours of defining a common housing cycle range from the identification of local house price booms and busts (Ferreira and Gyourko 2011) to clustering MSAs with similar cyclical patterns (Hernández-Murillo et al. 2017).

[5]We have also used the Harding and Pagan (2002) algorithm based on local minima and maxima to check the proportion of MSAs that share the same peak and trough as defined by the median. Results are broadly consistent with our approach.

[6]In a sample of 79 MSAs, Glaeser et al. (2008) identify a national boom over 1982-1989, a subsequent bust until 1996, and a strong boom between 1996 and 2006. We get a different picture for 1986-1996, since we cover a substantially larger sample of MSAs.

Figure 3: Real house price cycles



Sources: Bureau of Labor Statistics, Federal Housing Finance Agency, Moody's Analytics, and authors' calculations.
Notes: Real house prices refer to the FHFA house price index, a weighted, repeat-sales index, deflated by CPI. The index assumes the value of 100 in 1995q1. The solid red line represents the US aggregate index, the long-dashed blue line the median for the MSA distribution, the yellow line with markers the 10^{th} percentile, and the dashed green line the 90^{th} percentile. The vertical lines divide the sample period by phases of the housing cycle.

to the first decile, while they increased by 93 percent for the top decile (Table 1). During the 2006-2012 bust, house prices fell in all, but one, MSA. By the end of 2017, house prices had increased in more than 90 percent of the MSAs since the trough of 2012.

Table 1: Local house price cycles

	US	Median	p10	p25	p75	p90	N	>0
1996-2006	51.5	32.7	16.6	22.0	64.4	93.1	254	254
2006-2012	-28.0	-21.2	-46.0	-31.7	-14.3	-10.0	254	1
2012-2017	23.3	13.3	1.3	6.2	27.4	52.2	254	237

Notes: Cumulative changes in real house prices for different phases of the housing cycle. The first column refers to the national index, and the following columns show points in the distribution for the MSA sample. *N* is the number of MSAs, while *>0* counts the MSAs that recorded cumulative house price increases over each cycle.

4 Estimating housing supply elasticities in booms

4.1 Main specification

To estimate local housing supply elasticities across the two housing booms, we use a single-equation approach in the spirit of Green et al. (2005). The authors estimate time-invariant housing supply elasticities for a sample of 45 MSA over the 1979-1996 period, by regressing a proxy for the annual growth in the housing stock on lagged house price growth. We use building permits as our housing supply variable to capture the immediate reaction of builders to a change in house prices.[7] Given that building permits do not exhibit stochastic non-stationarities, we adopt a level specification. We follow Glaeser et al. (2008) and assume that permits depend on the price-to-cost ratio (Tobin's Q). Due to data availability, we use wages and salaries in the construction sector as a proxy for total construction costs. We account for geographical (Saiz 2010) and regulatory constraints (Gyourko et al. 2008) in the response of housing supply to a change in house prices. We estimate the following specification separately for the two boom periods:

$$\begin{aligned}\log(H_{i,t}) =& \beta^j \log(HPI_{i,t}) + \lambda^j \left[\log(HPI_{i,t}) \times UNAVAL_i\right] + \delta^j \left[\log(HPI_{i,t}) \times WRLURI_i\right] \\ &+ \gamma^j X'_{i,t} + \eta_i^j + \zeta_t^j + \epsilon_{i,t}^j\end{aligned} \tag{1}$$

where $log(H_{i,t})$ denotes the log of building permits, $\log(HPI_{i,t})$ is the log of the FHFA house price index deflated by CPI, $UNAVAL_i$ is the land unavailability index of Saiz (2010), $WRLURI_i$ is the Wharton Land Use Regulatory Index (Gyourko et al. 2008), and $X'_{i,t}$ is a vector of local economic and socio-demographic variables, which includes the lagged dependent variable, the log of real construction wages, and its interaction with the two supply restriction indices, log of population, the unemployment rate, the inflation rate, the dependency ratio, and the fraction of Blacks and Hispanics in total population. We add η_i^j to account for MSA-fixed effects, and ζ_t^j to capture time-fixed effects. The superscript j indicates that the estimated parameters may differ across the two booms, $j = \{1996-2006, 2012-2017\}$.

We expect β^j to be positive, as builders apply for more building permits when house prices increase. In addition, the interaction terms in Eq. (1) implies that housing supply elasticities may differ across MSAs if there are differences in land availability or regulation. We expect the coefficients λ^j and δ^j to be negative, as tighter geographical or regulatory restrictions should lead to a smaller expansion in building permits. It follows that the implied supply elasticity for

[7] The process of building a housing unit first requires builders to apply for a permit to get their construction project approved, which can take a few months. After the approval is granted, the construction works start (housing starts). The process ends when the housing unit is occupied or becomes available for occupancy (housing stock).

a given MSA in housing boom j is found by differentiating Eq. (1) with respect to house prices:

$$Elasticity_i^j = \beta^j + \lambda^j \times UNAVAL_i + \delta^j \times WRLURI_i \tag{2}$$

4.2 IV identification

To deal with reverse causality between house prices and permits, we use an IV approach. An instrument, Z, for house prices in the housing supply equation needs to shift housing demand (and thereby house prices), while at the same time be orthogonal to omitted supply factors. More formally, the traditional IV conditions for all i and t need to be satisfied:

$$Cov(Z_{i,t}, HPI_{i,t}) \neq 0 \tag{3}$$

$$Cov(Z_{i,t}, \epsilon_{i,t}) = 0 \tag{4}$$

where Eq. 3 is the *relevance* condition, stating that the external instrument Z must be contemporaneously correlated with local house prices. The *exogeneity* condition in Eq. 4 requires the instrument not to be contemporaneously correlated with the omitted supply factors in Eq. 1.

We use two instruments for house prices that we argue lead to shifts in housing demand (relevance), but that does not shift housing supply (exogeneity).[8] The first instrument exploits variation in crime rates across MSAs and over time. We use data on crime rates (per 100,000 inhabitants) from the Uniform Crime Report Offenses Known to Law Enforcement data set, which is compiled by the FBI. These data provide counts of crimes reported to the police for each police agency (cities, towns, and villages), and broken down by two major types: violent crime (murder, forcible rape, robbery, and aggravated assault), and property crime (burglary, larceny theft, and motor vehicle theft). Given the significant negative impact that crime can have on society, either directly through destruction of life and of property, or indirectly through the creation of a sense of insecurity, fear and anxiety as a consequence of criminal acts, crime can be viewed as a negative amenity (Pope and Pope 2012). Accordingly, crime rates should capture exogenous variation in (negative) amenities that drive house price changes both within and across MSAs.

[8] We cannot use supply shifters as instruments as they would not satisfy the orthogonality condition. In particular, we cannot resort to one of the most commonly used instruments for house prices, namely the housing supply elasticity calculated by Saiz (2010), see e.g., Mian et al. (2013), and Stroebel and Vavra (2019) – although not free of criticism (Davidoff 2016). The reason is that it enters the supply equation that we are interested in estimating, and because the housing supply elasticity is our main parameter of interest.

The relevance condition is supported by findings in the literature that point to high crime rates being strongly and negatively associated with property prices. The seminal paper by Thaler (1978) finds that an increase in property crime per capita reduces house prices in Rochester, New York. More recent papers have found a detrimental effect of crime on property prices, such as Gibbons (2004) for London. In turn, Schwartz et al. (2003) estimates that falling crime rates were responsible for one-third of the increase in property values in New York over 1994-98. Along the same lines, but using zip code-level data, Pope and Pope (2012) estimates the elasticity of property values to the decline in crime rates over 1990-2000 to have been important. We use property crime, which accounts for almost 90 percent of total crime, as our main measure of crime since it is available for a larger sample of MSAs compared with violent crime.[9]

The second instrument we use is the log of real personal disposable income. Income is one of the main determinants of housing and consumption demand in standard macro and housing models, but typically does not affect housing supply directly (Dougherty and Van Order 1982, Buckley and Ermisch 1983, Meen 1990, Muellbauer and Murphy 1997, Meen 2001, 2002, Duca et al. 2011). This instrument should thus satisfy both the relevance and exogeneity conditions.

The validity of the instruments hinges on property crime rates and income affecting housing supply *only* through its impact on house prices, i.e., leading to movements along, but not shifts in, the supply curve. One potential concern is that housing supply conditions may be endogenous to property crime, invalidating the use of our instrument. One could argue that less affordable housing may lead to more property crime, implying a negative association between crime and house prices. On the other hand, one could also argue that high-income neighborhoods are more prone to property crime, implying a positive association between crime and house prices. While these are admittedly possible concerns when using data at the granular level, they are less likely to be present when using MSA data as neighborhood (zip code) level effects are washed out in the aggregation.[10]

Although it is impossible to formally test the exclusion restriction, we provide some evidence that it is valid in our context. First, we minimize this bias by adding several local supply controls to the regression. Second, we examine the exclusion restriction along the lines of Mian and Sufi (2011). They use Saiz (2010)'s housing supply elasticities to instrument for house

[9] In Section 7 we show that none of our results are materially affected by instead using total crime as the instrument.

[10] Note also that the MSA fixed-effects in our panel model should capture the potential time-invariant endogeneity between supply conditions and MSA-idiosyncratic characteristics.

prices, and validate their exclusions restriction by showing that wage growth did not accelerate differentially in elastic and inelastic areas over the 2002-2006 period. Table 2 shows that crime rates and income are not associated with statistically different wage growth developments in the construction sector in any of the two booms. The same holds true when the dependent variable is the level of construction wages, so as to allow for the possibility that crime rates can also have a permanent level effect on wages.

Table 2: Validity of the exclusion restriction

Dep. var:	1996-2006		2012-17	
Wage growth	(1)	(2)	(3)	(4)
$\log(Crime)$	-0.079	0.043	-0.045	-0.014
	(0.068)	(0.081)	(0.124)	(0.150)
$\log(Inc)$	-0.014	0.310	-0.022*	-0.412
	(0.013)	(0.285)	(0.011)	(0.247)
Controls	No	Yes	No	Yes
Number of MSA	242	241	254	254
Observations	7,584	7,548	4,866	4,866
Adj. R2	0.439	0.446	0.263	0.263

Notes: OLS estimates with state-fixed effects and time effects, where the dependent variable is the change in the log of construction wages. The constant and control variables are not reported. Robust heteroskedastic standard errors shown in parentheses. Asterisks, *, **, and ***, denote statistical significance at the 10%, 5%, and 1% levels.

We have *three* endogenous regressors, as house prices interacted with the supply restrictions *UNAVAL* and with *WRLURI* are also endogenous. We therefore have *six* instruments. For each boom, we estimate the following first- and second-stage regressions:

$$\begin{aligned} W_{i,t} =& \rho_1^j \log(Crime_{i,t}) + \rho_2^j [\log(Crime_{i,t}) \times UNAVAL_i] + \rho_3^j [\log(Crime_{i,t}) \times WRLURI_i] \\ &+ \omega_1^j \log(Inc_{i,t}) + \omega_2^j [\log(Inc_{i,t}) \times UNAVAL_i] + \omega_3^j [\log(Inc_{i,t}) \times WRLURI_i] \\ &+ \phi^j X'_{i,t} + \psi_i^j + \nu_t^j + \mu_{i,t}^j \end{aligned} \tag{5}$$

$$\begin{aligned} \log(H_{i,t}) =& \beta^{IV,j} \log(\widehat{HPI_{i,t}}) + \lambda^{IV,j} \widehat{[\log(HPI_{i,t}) \times UNAVAL_i]} + \delta^{IV,j} \widehat{[\log(HPI_{i,t}) \times WRLURI_i]} \\ &+ \gamma^j X'_{i,t} + \eta_i^j + \zeta_t^j + \epsilon_{i,t}^j \end{aligned} \tag{6}$$

where j signifies again that all parameters may differ between the two booms. The dependent variable, $W_{i,t} = \{HPI_{i,t}, HPI_{i,t} \times UNAVAL, HPI_{i,t} \times WRLURI\}$ in Eq. 5 refers to house prices, and house prices interacted with supply restrictions. To control for possible confounders, we add a set of control variables, listed in Section 4.1. We assess the relevance and strength of the instruments with the weak identification Cragg-Donald F-statistic test, including a version of the test that is robust to heteroskedasticity (Kleibergen-Paap F-test.) We take Stock and Yogo (2005)'s critical value of 12.2 for the 5 percent relative bias to test for weak instruments. We also compute the Hansen J-statistic test to test for over-identification, given that we have more instruments than endogenous variables.

Results are reported in Table 3 for both the 1996-2006 boom and the 2012-2017 boom. The first-stage F-test and robust F-test stand between 30 and 50, which is significantly above Stock and Yogo (2005)'s threshold value, suggesting that our instruments are valid and strong.[11] In addition, the Hansen J-test provides strong evidence against rejecting the null hypothesis that the instruments are valid in the first boom. We reach a similar conclusion for the second boom, although the evidence is somewhat weaker. The coefficient on house prices is statistically significant at conventional levels, and positive, for both housing booms. But there is a considerable decline in the magnitude of the coefficient from the first to the second boom. This implies a weakened response of permits to a given change in house prices. Our estimates indicate that building permits increased by 2.8 percent over the short term (long-term response of 4.7 percent) for every 1 percent increase in house prices during the 1996-2006 boom, which is almost twice as large as during the current housing recovery – a response of roughly 1.8 percent over the short term (long-term response of 2.2 percent).[12]

The interaction of house prices with the supply restriction variables yields the expected negative signs, i.e., the tighter the geographical and regulatory restrictions, the smaller is the expansion in building permits for a given house price increase. The coefficient on the interaction term for $UNAVAL$ is, however, not significant in the current boom.

[11] The first-stage coefficients on the instruments are statistically significant for both housing booms: for property crime rates we get coefficients within a range of -0.02 to -0.025 (t-stats above 2), and of around 0.3-0.4 (t-stats above 8) for income.

[12] The long-term coefficient is the result of dividing its short-run coefficient by 1 minus the lagged coefficient on the dependent variable; for instance, for the 1996-2006 cycle: 4.7=2.774/(1-0.415).

Table 3: Regression estimates by housing boom

	1996-2006	**2012-2017**
$\log(HPI)$	2.774***	1.794**
	(0.428)	(0.847)
$\log(HPI) \times UNAVAL$	-1.344***	-1.225
	(0.340)	(1.185)
$\log(HPI) \times WRLURI$	-0.718***	-1.086**
	(0.096)	(0.422)
$\log(H_{t-1})$	0.415***	0.203***
	(0.019)	(0.023)
Number of MSA	241	254
Observations	7,548	4,866
Cragg-Donald F-test	39.83	49.66
Kleibergen-Paap (robust) F-test	31.00	29.61
Hansen J-test (p-value)	0.64	0.06

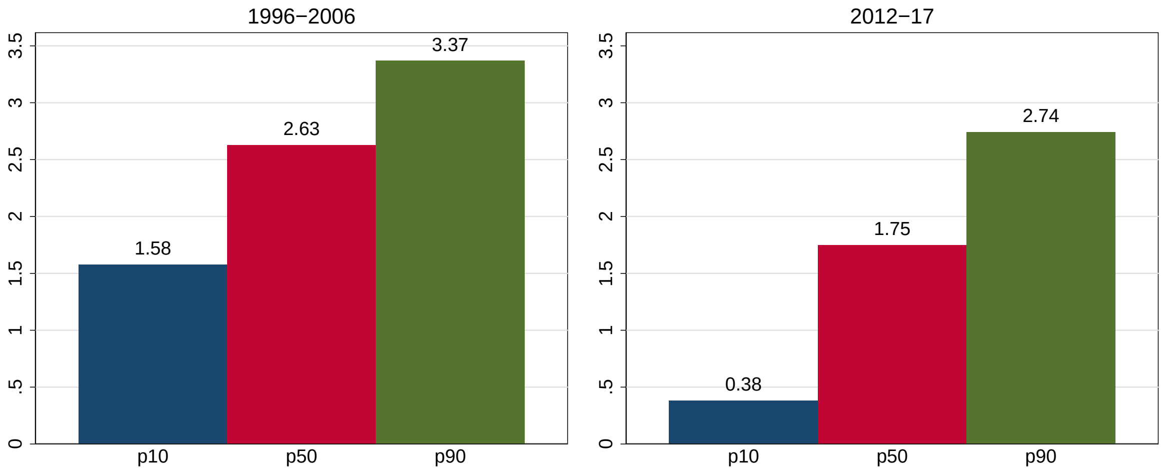
Notes: IV estimates of Eq. 6, where the dependent variable is the log of building permits. The Cragg-Donald F-test and Kleibergen-Paap F-test assume that under the null the excluded instruments are not weakly correlated with the endogenous regressors. The Hansen J-test of overidentifying restrictions reports the p-value under the null hypothesis that the instruments are uncorrelated with the error term, and that the excluded instruments are correctly excluded from the estimated equation. The constant and additional control variables are not reported. Robust heteroskedastic standard errors shown in parentheses. Asterisks, *, **, and ***, denote statistical significance at the 10%, 5%, and 1% levels.

4.3 Estimated elasticities

We calculate MSA-specific elasticities for the two booms by inserting the relevant parameters of Eq. 6 into the expression of Eq. 2. Figure 4 shows the elasticities at the median, 10^{th} and 90^{th} percentiles for each housing boom. Our results suggest that supply elasticities have fallen across the whole distribution. In addition, the dispersion in supply elasticities has increased during the current cycle, with a particularly strong decline in the lowest part of the distribution.

We shed more light on the heterogeneity between MSAs by looking at the distribution of the elasticities across the two housing booms (Figure 5). More specifically, we create five groups, where red (blue) colors refer to low (high) elasticity areas. Areas located in states such as California, Arizona, Florida, Oregon, and New York have the lowest elasticities in both booms. This is not surprising, given that geographical idiosyncrasies, such as steep ground and bodies of water, make it harder to build and limit the land available for construction in these areas, compared to the rest of the country (Saiz 2010). In addition, land-use regulation, which limits the expansion of supply, also tends to be more stringent in these areas (Gyourko et al. 2008). By contrast, we estimate high-elasticity areas to be located in the Midwest, where builders face relatively fewer restrictions to expand housing supply.

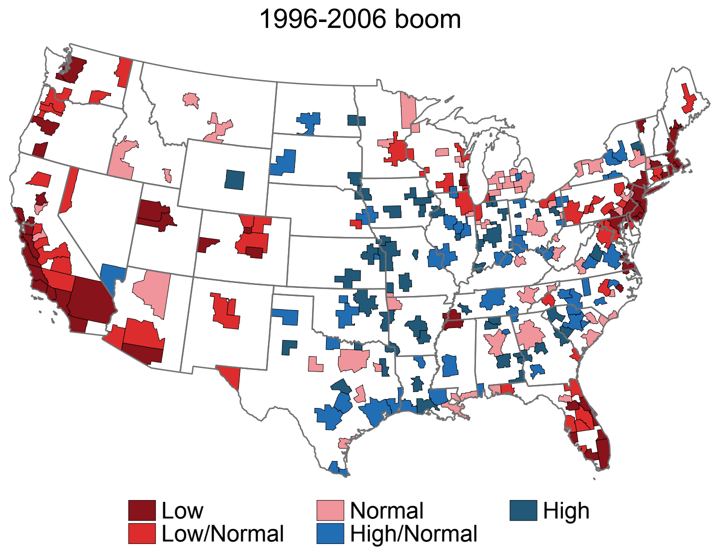
Figure 4: Estimated elasticities: IV specification



Notes: Estimated elasticities from Eq. 6 for the median, 10^{th} and 90^{th} percentiles for each housing boom.

Figure 5 shows that the rank ordering of the MSAs between the two booms is relatively stable, and Figure 6 reveals that the largest decline in elasticities between the two booms has taken place in the areas with the lowest elasticities during the first housing boom.

Figure 5: Estimated elasticities for the two housing booms



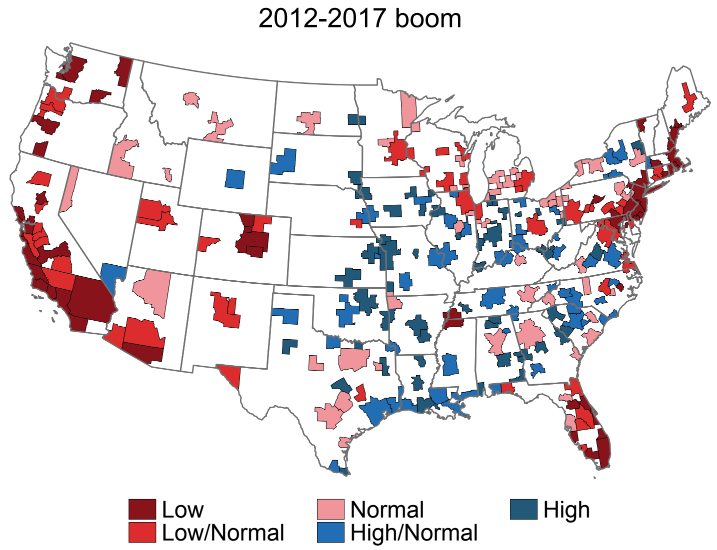
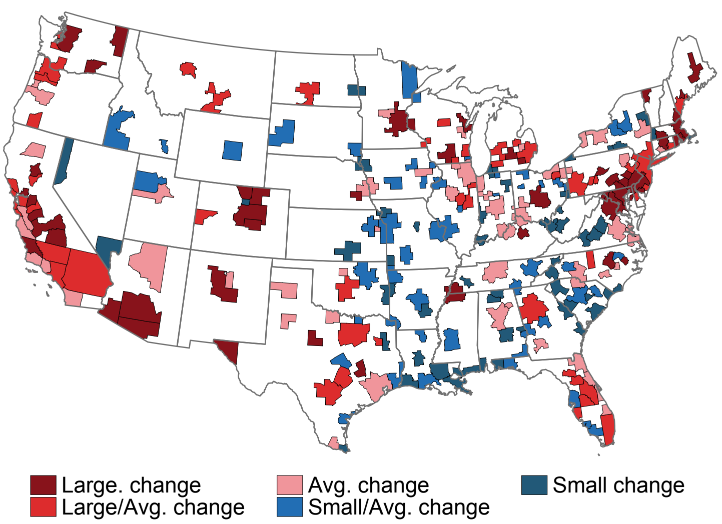


Figure 6: Change in estimated elasticities between booms



5 Supply elasticities and demand shocks across booms

Our results point to a nationwide decline in housing supply elasticities. An implication of this is that aggregate demand shocks should have a greater impact on house prices today, whereas quantity should respond less (see Appendix A for the illustration of this point in a simple supply-demand framework). We explore the relevance of this conjecture through the use of exogenous monetary policy shocks.

5.1 High-frequency identification of monetary policy shocks

Our measure of monetary policy shocks is computed following a recent strand of the literature that resorts to high-frequency data to identify unexpected changes in the Fed policy rate (see, for instance, Gürkaynak et al. 2005, Gertler and Karadi 2015, Nakamura and Steinsson 2018).[13] This high-frequency identified (HFI) approach isolates news about future policy actions that is orthogonal to changes in economic and financial variables. We take the unexpected changes in interest rates for 3-month ahead contracts on Fed funds futures in a 30-minute window surrounding FOMC meetings. In total, we cover 127 meetings over the two housing booms: 83 between 1997q1-2006q4 and 44 between 2012q3-2017q4. The underlying assumption is that changes in the futures rates within that window can only arise from news about monetary policy, given that market participants incorporate all publicly available information at the beginning of that narrow window.

More specifically, let f_{t+j} be the price of a Fed funds future in month t that expires in j months, and S_{t+j} the unanticipated change in the expectation for the Fed funds rate $t+j$ months ahead. The monetary surprise is then constructed as the difference between the price of the $t+j$ month ahead Fed funds future contract 20 minutes after the FOMC announcement and the price of the same contract 10 minutes before the announcement:

$$S_{t+j} = f_{t+j} - f_{t+j,-1}$$

We follow standard practice in transforming high frequency data into the quarterly frequency (see, for instance, Ottonello and Winberry 2018, Wong 2019). In particular, we first create a daily shock series by cumulating the daily surprises over the past 90 days. We then take quarterly

[13] We do not use the standard Romer and Romer (2004)'s narrative shocks given that the Greenbook projections are not available for the period covering the current recovery; they are released to the public with a lag of five years.

averages of the cumulative daily shocks. Our quarterly shocks are characterized by roughly a 60-40 distribution between expansionary and contractionary shocks over the full sample (Figure D.5 in Appendix D).[14]

HFI shocks may contain measurement errors, thus may capture only part of the 'true' structural shock. For instance, some price changes within the 30-minute window around the policy announcements may reflect trading noise and volatility. In addition, the monthly (and quarterly) series of surprises contains some random zero observations, as a result of calendar months without FOMC meetings. Finally, the monthly (and quarterly) surprise series does not incorporate other monetary policy news released outside of the announcement window, such as speeches by FOMC members. To deal with this, we follow Gertler and Karadi (2015), Ramey (2016), Nakamura and Steinsson (2018), Stock and Watson (2018) and treat the surprises as instruments for the underlying shock. Following Gertler and Karadi (2015), we choose the one-year Treasury bill yield as the relevant monetary policy indicator. This risk-free asset with a longer maturity than the funds rate has the advantage of also incorporating shocks to forward guidance about the future path of interest rates, instead of just about the current rate.

5.2 Empirical results: LP-IV

To study how monetary policy shocks affect house prices and quantity across MSAs over the two booms, we follow Jordà et al. (2015), Ramey (2016), and Stock and Watson (2018) and use an instrumental variable local projection approach. The Jordà (2005) method offers some advantages over Vector Auto Regressive (VAR) models, since impulse responses are less vulnerable to mis-specification (Stock and Watson 2018). In addition, it easily accommodates non-linearities, allowing us to estimate the dynamic causal effects of monetary policy shocks conditional on our housing supply elasticities. We estimate the LP-IV model over one unique sample, the two booms 1997q1-2006q4 and 2012q3-2017q4, by running a series of regressions for each horizon h=1,2...,16 quarters:

$$\Delta_h Y_{i,t+h} = \beta^{Y,h} \Delta MP_t + \gamma^{Y,h} \Delta MP_t \times \widehat{Elast}_i^j + \sum_{j=1}^{4} \lambda_j^{Y,h} \Delta X_{i,t-j} + \eta_i^{Y,h} + \epsilon_{i,t+h}^{Y} \tag{7}$$

where the dependent variables, Y, are the cumulative percentage change in real house prices,

[14]The time-aggregation bias should not affect the results, as our quarterly shocks exhibit similar moments to the raw high-frequency data (Table D.1 in Appendix D).

HPI, or in building permits, H, from period t to $t+h$.[15] MP_t is the monetary policy indicator (the one-year Treasury bill yield), which is interacted with our estimated supply elasticities $\hat{Elast}_{i,t}$ for each boom, and $X_{i,t-j}$ refers to a vector of lagged control variables (four lags), namely the lagged dependent variables, the external instrument, real disposable income growth, population growth, real construction wage growth, the change in the unemployment rate, and the Gilchrist and Zakrajšek (2012)'s excess bond premium (EBP).[16] This large set of control variables helps minimize the omitted variable bias and reduce the variance of the error term (Stock and Watson 2018). In addition, Stock and Watson (2018) argue that the nature of the construction of the HFI monetary shocks induces a first-order moving average structure, leading to a correlation between the external instrument and past values of the policy indicator. We follow their suggestion and include lagged values of the external instrument as controls to make our IV valid.

We add MSA-fixed effects $\eta_i^{Y,h}$ to control for time-invariant idiosyncratic MSA characteristics, but we do not include time-fixed effects given that the monetary policy indicator is common across MSAs. The standard errors are MSA-specific cluster-robust, which allow for fully flexible time dependence in the errors within MSAs.[17]

Our parameters of interest are $\beta^{Y,h}$ and $\gamma^{Y,h}$. Following the conjectures from the theoretical model in Appendix A, we expect an expansionary monetary policy shock to boost house prices ($-\beta^{HPI,h} >0$), but that this effect becomes smaller the higher the housing supply elasticity ($-\gamma^{HPI,h} <0$). Further, we expect an expansionary shock to stimulate more construction activity ($-\beta^{H,h} >0$), and that this effect is reinforced by a higher elasticity ($-\gamma >0$).

We have two endogenous variables and two instruments in Eq. 7: the monetary policy indicator and its interaction with the estimated elasticities, instrumented with the HFI surprise series and with its interaction with the elasticities. The first-stage F-test and robust F-test are above the Stock and Yogo (2005)'s threshold, suggesting that our instruments are valid and strong. We find that an expansionary monetary policy shock that lowers the one-year Treasury

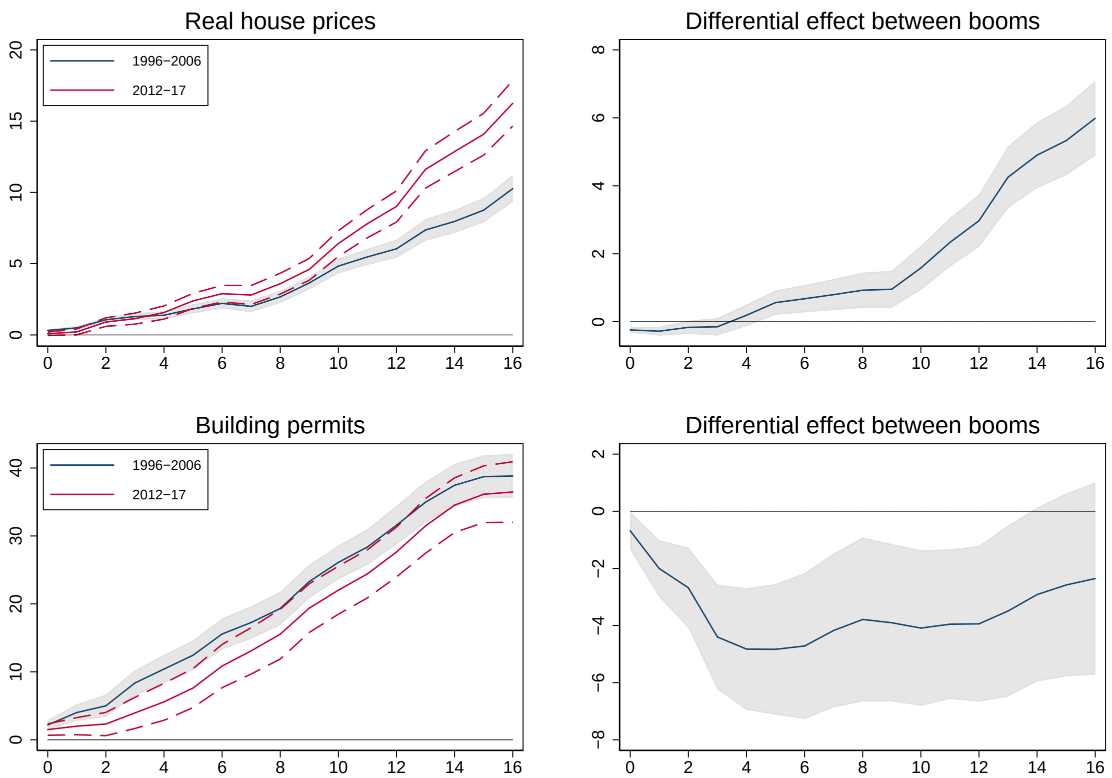
[15]Given the high volatility of permits, especially as h increases, we transform the raw series into a four-quarter centered moving average.

[16]The EBP is a measure of investor sentiment or risk appetite in the corporate bond market that is not directly attributable to expected default risk. More specifically, Gilchrist and Zakrajšek (2012) define it as the spread between the rate of return on corporate securities and a similar maturity government bond rate that is left after removing the default risk component. We add the EBP as Gertler and Karadi (2015) argue that it has strong forecasting ability for economic activity, thus acting as a summary indicator of the potentially relevant information left out of the model to explain the dependent variable.

[17]This adjustment tends to produce more conservative standard errors than a standard heteroskedasticity-and-autocorrelation (HAC) estimator (Jordà et al. 2015). Note that the standard errors are not distorted by the generated regressor issues, given that the high-frequency shock is used only as an instrument and not directly included in the model.

bill yield by 100 basis points raises both house prices and quantity over the short to medium run in a statistically significant way for both housing booms (Figure 7). Furthermore, we find that house prices rise by considerably more in the 2012-2017 boom compared with the 1996-2006 boom. While price dynamics are similar in the short term, house prices in the current boom start to increase at a statistically significant faster pace after two years. For the same 100 basis points decline in government bond yields, real house prices in the current boom are six percentage points higher after four years (a cumulative 16 percent increase in the 2012-17 boom against ten percent in the previous boom). We estimate the opposite dynamics for building

Figure 7: Responses to an expansionary monetary policy shock across booms



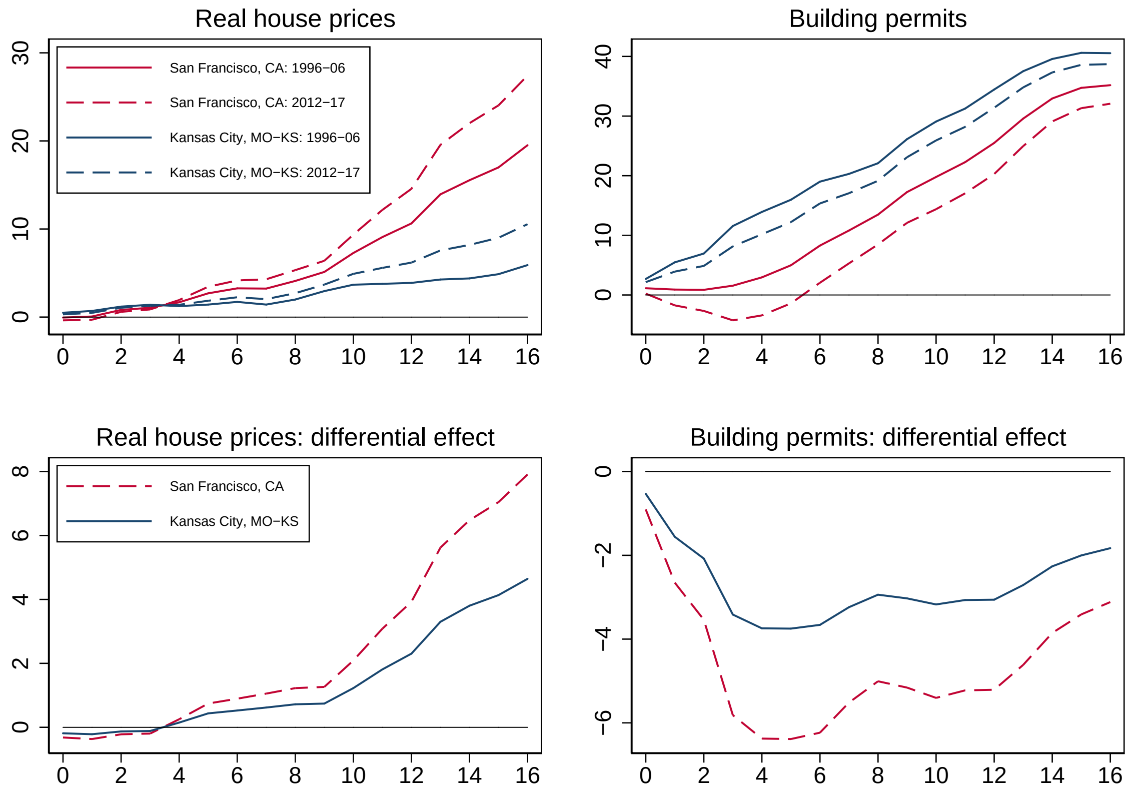
Notes: Cumulative impulse responses to a 100 basis point decline in the one-year Treasury bill yield, assessed at the sample median elasticity for each housing boom period. The right-hand charts depict the difference in the estimated response of house prices and building permits between the 2012-17 and the 1996-2006 booms. The grey area and the dashed red lines refer to 90 percent confidence bands.

permits, which reacted more strongly to a monetary policy shock in the 1996-2006 boom. But the difference between the responses is relatively small, given the scale of the increase in permits in both episodes (almost 40 percent after four years). Overall, the differences in the impulse responses are not driven by different magnitudes of the underlying shocks, as illustrated by a similar decline in the response of the policy indicator (Figure D.6 in Appendix D).[18]

[18] We also check that the statistical difference in the impulse responses between the two booms are robust to adjusting the standard errors for cross-sectional dependence using the Driscoll-Kraay estimator (Figure D.7). Our results are broadly robust to this.

We show that there is considerable heterogeneity in responses across MSAs within the same period. Figure 8 shows that house prices in a typical low-elasticity MSA, such as San Francisco-Oakland-Hayward, California respond more strongly to the monetary policy shock than a typical high-elasticity MSA, such as Kansas City, Missouri. While this is in line with Aladangady (2014) and Aastveit and Anundsen (2017), our results also suggest that the differential effect between the two booms may be larger in low-elasticity areas than in high-elasticity areas (lower panel of Figure 8). Although it is outside the scope of this paper, the time-varying effects of monetary policy also raise concerns about the distributional effects of monetary policy on consumption inequality between MSAs (Beraja et al. 2019).

Figure 8: Responses to an expansionary monetary policy shock for selected MSAs



Notes: Cumulative impulse responses to a 100 basis point decline in the one-year Treasury bill yield, assessed at the sample median elasticity for selected MSAs and for each housing boom. Kansas City, Missouri, represents a high-supply elasticity MSA, while San Francisco-Oakland-Hayward, California, a low-supply elasticity MSA.

6 Why have elasticities declined?

In theory, several factors might lead to changes in the slope of the housing supply curve, including changes in regulatory conditions, demographics, and in expectations about future demand and house prices. A recent paper by Herkenhoff et al. (2018) documents a substantial tightening in land-use policy in most US states since 1950. They find that a substantial tightening across states took place between 1990 and 2014, of around 18 percent. The tightening in regulation is particularly marked for high-house price states. Along the same lines, recent research has put forward the notion that the decline in construction productivity may be the result of increased costs stemming from tighter regulation over time (Davis and Palumbo 2008, Albouy and Ehrlich 2018) and Glaeser and Gyourko (2018).

A simple correlation analysis between our estimated elasticities and Herkenhoff et al. (2018)'s land-use regulation index suggests that the tightening in regulation between 2000 and 2014 is associated with a decline in our estimated elasticities between the two housing boom episodes (correlation of -0.4).[19] We show that this relationship holds in a multi-variate setting, by estimating the following cross-sectional equation:

$$\Delta \log(\widehat{Elast}_i^{17} - \widehat{Elast}_i^{06}) = \alpha + \beta_1 \Delta \log(X_i^{17} - X_i^{06}) + \beta_2 Z_i + \epsilon_i \tag{8}$$

where the dependent variable is the log percentage change in estimated elasticities between 2012-2017 ($\widehat{Elast}_i^{17}$) and 1996-2006 ($\widehat{Elast}_i^{06}$). We regress it on the log percentage change for the same period of a set of indicators X_i, namely the state-level Herkenhoff et al. (2018)'s land-use regulation, population density, construction wages, unemployment rate, and on initial conditions Z_i, the levels of house prices to income per capita and of population density. We also include the cumulative change in house price growth during the 2006-2012 bust.

Our results provide statistical evidence that tighter land-use regulation has been associated with a decline in elasticities between the two booms (Table 4).[20] Our estimates also show that areas with stronger economic performance, as measured by the change in the unemployment rate, and higher initial levels of house prices relative to income and of population density at the beginning of the 2012-2017 boom, tend to be associated with larger declines in elasticities. In contrast, the negative association between faster population density growth and larger declines in elasticities is not statistically significant.

Finally, we find that the areas that experienced the strongest bust in house prices over the period 2006-2012 ($\Delta \text{HPI}^{06-12}$) also recorded the largest declines in elasticities between the two booms. Our interpretation is that the Great Recession might have cast a long shadow on

[19] Herkenhoff et al. (2018)'s land-use regulation indicator is available for 48 states, excluding Alaska and Hawaii, and for individual years: 1950, 1960, 1970, 1980, 1990, 2000, and 2014. We take the 2000 and 2014 values of that indicator as the data points relevant for respectively the 1996-2006 and 2012-2017 booms.

[20] A decline in the land-use regulation index represents a tightening in regulation.

builders' expectations, making them less price responsive than before. This fear of a new bust may have paved the way for a new housing boom where house prices are more responsive to fluctuations in demand.

Table 4: ΔElasticity between booms

	(1)	(2)	(3)
$\Delta\log$(Land reg.)	0.273***	0.249***	0.162***
	(0.043)	(0.038)	(0.043)
ΔHPI^{06-12}		0.886***	0.935***
		(0.095)	(0.150)
$\Delta\log$(Pop density)			-0.008
			(0.138)
$\Delta\log(Wage)$			-0.010
			(0.084)
ΔUR			3.515**
			(1.558)
Hpinc_pc			-0.568***
			(0.139)
Pop density			-0.011**
			(0.004)
Observations	251	251	251
R-squared	0.121	0.379	0.465

Notes: Regression estimates of Eq. 8, where the dependent variable is the percentage change in the estimated supply elasticities between 2012-2017 and 1996-2006. Robust heteroskedastic standard errors in parentheses. Asterisks, *, **, and ***, denote statistical significance at the 10%, 5%, and 1% levels.

7 Robustness checks and alternative explanations for declining supply elasticities

7.1 Bartik-type instrumental variable approach

We check the robustness of our baseline estimates of housing supply elasticities to employing a Bartik-type instrumental variable approach (Bartik 1991). More specifically, we follow a similar approach as Guren et al. (2018), and instrument MSA-level house prices with house prices at the Census Division level.[21,22] A detailed description of the approach is provided in Appendix

[21]The nine Census Divisions are New England, Middle Atlantic, East North Central, West North Central, South Atlantic, East South Central, West South Central, Mountain, and Pacific.

[22]This is akin to a Bartik-type instrument, as the strategy employed assumes that house prices in a given number of MSAs respond differently to an aggregate shock (regional house price changes) because of pre-existing local differences in the housing market or economic structure. In the original setting, the Bartik instrument involves instrumenting local employment growth with a variable that consists of the interaction between local

C. Our results are broadly robust to this approach. The estimated elasticities are in line with our baseline results, with a larger decline in the elasticities in the current boom, see Figure D.8 in Appendix D.

7.2 Alternative specifications for estimating housing supply elasticities

We carry out additional robustness checks to the housing supply equation (6) by: (i) using total crime rates (sum of property crime and violent crime) as the crime variable instrument; (ii) using permit intensity as the dependent variable to allow the dynamics in permits to differ according to the existing stock of houses; (iii) replacing UNAVAL and WRLURI with a summary measure of supply restrictions, essentially the sum of these two variables standardized, to account for the possibility that these indicators might be correlated; and (iv) controlling for mortgage originations (the amount of new mortgage loans) to assess the impact on the housing supply response of subdued credit developments since the Great Recession.[23,24] Our results are robust to these alternative specifications (Table D.3 in Appendix D). Moreover, across all specifications, our finding that supply elasticities have declined between the two housing booms is maintained (Table D.4).

7.3 Rolling window estimation of housing supply elasticities

Our approach has been to estimate housing supply elasticities for the two boom periods separately. Another approach is to estimate housing supply elasticities using a rolling window estimation. To explore whether this has any bearing on our findings of a decline in housing supply elasticities, we estimate Eq. 6 using 10-year and 15-year rolling windows. For the 10-year window, the first regression covers the period 1997–2006, the second regression spans the period

industry employment shares and national industry employment growth.

[23] Additional robustness checks we have carried out include by: (i) using housing starts and the housing stock as the dependent variables; (ii) using the Arellano-Bond estimator to account for the Nickell (1981) bias in dynamic panels; (iii) adding state-by-time fixed effects to control for time-varying state-specific shocks; and (iv) by estimating the supply equation separately for multi-family building permits. Our baseline regression estimates remain qualitatively unchanged. Furthermore, in a previous version of the paper, we used the mean January temperature instead of crime rate as one of the instruments for house prices, based on the work of Glaeser and Gottlieb (2009), and Glaeser et al. (2012). January temperatures proxy housing demand as they capture the exogenous variation in amenities that lead house prices to change. We find qualitatively similar results as the baseline specification used throughout this paper. The drawback, however, is that the mean January temperature turns out to be a weaker instrument for house prices as it is only able to identify house prices in the cross-section; given the small variability over time, it is defined as monthly average temperatures in January calculated over 1941-1970. Details are available upon request.

[24] Although we would ideally like to control for changes in credit conditions of home builders, which can lead to a shift in the supply curve, data on credit to construction firms are not available at the MSA level. We use instead mortgage originations which should be correlated with the dynamics in credit to construction firms.

1998–2007, and so on. Similarly, for the 15-year rolling windows, the first period goes from 1997 to 2011, and the last from 2003 to 2017. We report the rolling window estimates of the median housing supply elasticity in Figure D.9. We find that housing supply elasticities have declined over time, in line with our baseline results. The durability of housing entails that housing supply is rigid downwards (Glaeser and Gyourko 2005), implying that housing supply elasticities should fall towards zero during severe busts. Consistent with this, our rolling window estimates show a particularly strong decline in housing supply elasticities during the recent housing bust.

7.4 Alternative specifications for the impact of monetary policy shocks

We tested the robustness of the local projection regressions of Eq. 7 to (i) using surprises in the two-month ahead Fed funds futures to compute the high-frequency monetary shock; (ii) taking the two-year Treasury note rates as the policy indicator; and (iii) running the main model with only one lag. Figure D.10 in Appendix D shows that our main results remain qualitatively robust, with the model that employs the two-year Treasury note rates as the policy indicator (*GS2*) displaying the strongest responses: irrespective of the specification used, house prices rise by considerably more in the 2012-2017 boom, at the expense of a slightly weaker supply response.

7.5 Alternative explanations for declining supply elasticities

A first alternative explanation centers on the strong rise in construction activity during the 1996-2006 boom that led to an oversupply of houses in the subsequent period, implying that there may be less need for new homes to be built. We do not, however, find support for the oversupply hypothesis. The housing stock per capita has been trending consistently downwards during the recovery period, whereas it increased over 1996-2006 (Figure D.11 in Appendix D). At the same time, the number of homes available for sale per capita are low. In turn, housing vacancy rates have shown similar developments across the two booms, with the rental vacancy rates actually going somewhat below pre-crisis levels (Figure D.12). Moreover, although the number of new foreclosures were higher at the beginning of the current boom, they started converging steadily to the levels seen before, whilst the months' supply of houses is only slightly above the levels recorded during the previous boom (Figure D.13).[25] Based on these indicators,

[25] The months' supply of houses measures the ratio of houses for sale to houses sold. It indicates how long the current for-sale inventory would last given the current sales rate if no new houses were built. It is a commonly used indicator to assess the strength of the housing market.

there seems to be little evidence of a supply overhang in the current recovery, suggesting that it cannot explain the low construction activity in the face of strong price developments.

Second, the weak response of builders during the current recovery could also be explained by difficulties in expanding capacity given the shortage of workers in the construction sector. The Joint Center for Housing Studies at Harvard University reports that, a large fraction of home builders cites the shortage of skilled workers as a significant problem (JCHS 2018). In addition, job openings and employment growth also appear to remain subdued in this sector (Rappaport 2018). Nevertheless, we find mixed evidence in favor of this being an explanation for the decline in elasticities. While the share of workers employed in the construction sector remains slightly lower than before the crisis, it has increased since the recovery, and approached pre-crisis levels (left panel of Figure D.14 in Appendix D). From a different perspective, employment in the construction sector is actually above pre-crisis levels; the number of construction workers necessary to build a house is larger than previously, a similar point made by Leamer (2015) – right panel of Figure D.14.

Third, land appears to have become scarcer, which limits the expansion in housing supply. There is some evidence that the number of vacant lots declined between 2008 and 2017, which is particularly more pronounced in the Western metros of San Francisco, San Diego, Seattle, Los Angeles, and Las Vegas (JCHS 2018). Furthermore, Rappaport (2018) reasons that there is limited availability of undeveloped land in desirable locations, as the outward expansion in supply towards the periphery in many metro areas may have reached its geographical limit, in a context where households are reluctant to take on increasingly long and congested commutes. Related to this, inadequate transportation spending may affect the substitutability between homes in the outskirts and more central locations (Green et al. 2005). The combination of a growing population and inadequate infrastructure spending may have resulted in a lengthening of commute times, leading to a steepening of the land price gradient.

Fourth, following the implementation of Basel III under the Dodd-Frank Act, US regulators have applied more stringent regulatory capital requirements on loans extended to construction and land development. While the Dodd-Frank Act effectively raised capital requirements from 8 percent to about 10-11 percent for C&I loans more generally, it raised required capital to 15 percent for loans to construction and land development. The stricter capital requirements may have contributed to shortages of buildable lots across the country, and consequently to a decline in housing supply elasticities.

A final explanation for the disconnect between house price developments and construction activity in recent years is related to increased market concentration in the home building sector. Haughwout et al. (2012) document that the market share of a few large firms started to increase rapidly in the run-up to the Great Recession. More recently, Cosman and Quintero (2019) also show that there has been a decline in the competitive intensity over the last decade among developers in the United States.[26] In a more concentrated market, firms can time their housing production to maximize profits without fear of pre-emption. Cosman and Quintero (2019) find that this has led to greater price volatility, less production, and fewer vacant unsold units, as firms with market power can decide to build when demand growth is strongest and charge prices higher above their marginal cost of production. This phenomenon is consistent with our finding of a nationwide decline in housing supply elasticities.

8 Conclusion

We have provided evidence of a substantial and synchronized decline in local housing supply elasticities from the 1996-2006 housing boom to the ongoing recovery that started in mid-2012. An implication of this finding is that the house price responsiveness to a given demand shock should be higher today, at the expense of a smaller increase in quantity.

When we estimate the effect of an exogenous monetary policy shock on house prices in each of the two booms, we have found that monetary policy has a substantially greater impact on house prices during the current recovery than during the previous boom. In contrast, we have found that the expansion in building permits is slightly smaller today. Furthermore, our results point to significant heterogeneity in the responses across local housing markets. In particular, we estimate a substantially larger response of house prices to a monetary policy shock in supply inelastic markets than in areas with an elastic supply.

Our findings suggest that the decline in supply elasticities has been largest in areas where regulation has tightened the most. We also find that supply elasticities have declined more in areas that experienced the largest bust in house prices during the Great Recession. We interpret this as evidence that the fear of a new bust has led developers to be less price-responsive than

[26] Cosman and Quintero (2019) argue that increased market concentration has been the result of three main factors: (i) several construction companies filed for bankruptcy in the wake of the 2007-2009 Great Recession; (ii) a federal legislative stimulus measure in 2009 that increased the ability of home builders to use previous years' losses to reduce their tax payments, which was highly beneficial to the largest companies; and (iii) an increase in the numbers of mergers, leading to a concentration of production in a smaller number of firms.

before. This behavior may have paved the way for a new housing boom where house prices are more responsive to fluctuations in demand.

The lowering of housing supply elasticities may explain why recent research finds that monetary policy has become more effective for financial variables; an aggregate shock that raises housing demand is absorbed mostly by price adjustments, rather than quantity adjustments. This finding can be important for financial stability considerations, whereby the actions of policy makers aimed at stimulating (housing) demand may have unintended effects by exacerbating the rise in house prices. In the current environment of tighter regulation and declining elasticities, our findings cast some doubts about the view that the recent housing market recovery looks 'healthier' and more sustainable compared to the previous boom.

Another implication of our findings relates to wealth inequality, particularly intergenerational inequality. The combination of high house prices and a tight supply of homes makes it difficult for young people and households with little liquid assets to become homeowners. This may have a direct impact on household inequality, by favoring existing homeowners, which tend to be older and wealthier, as their housing equity increases. Despite the recent findings in the literature about the economic costs of regulation, local zoning laws have actually been tightening across the country, and this has reduced supply elasticities. The biggest challenge in relaxing local housing restrictions comes from existing homeowners not wanting more affordable homes, as higher house prices mean that the value of their asset goes up. In addition, existing homeowners also want to protect the amenities in their city, as new housing brings in more people, creating a congestion in access to public goods, such as crowded schools and roads (Glaeser and Gyourko 2018).

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Appendix A: Theoretical framework

We take as a starting point a simple supply-demand model with durable housing inspired by Glaeser et al. (2008), which is made up of an economy that contains a collection of several local housing markets that exhibit heterogeneity in economic, financial, and social dimensions, including in the supply elasticities. Abstracting from depreciation of the existing stock, the law of motion of capital accumulation for each area i in each period t is given by:

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

where $I_{i,t}$ is new investments in housing capital. We assume that the marginal cost of construction $MC_{i,t}$ is inversely proportional to the existing housing stock $H_{i,t-1}$, implicitly meaning that investment in new construction projects is more attractive in larger housing markets:

$$MC_{i,t} = C_{i,t} \times \left(\frac{I_{i,t}}{H_{i,t-1}} + 1\right)^{\frac{1}{\varphi_{i,t}}}$$

where $C_{i,t}$ represents housing construction costs (land, labor, and building materials), which rise with investment to reflect the scarcity of the inputs used into housing production, and $\varphi_{i,t}$ is the local-specific housing supply elasticity that is allowed to vary over time. The assumption of a time-varying supply elasticity, consistent with our estimates in the previous section, is a new feature of our model compared with Glaeser et al. (2008). We apply Tobin's Q theory to determine new investments, in that builders adjust supply based on the price of housing relative to the marginal cost of construction. The investment function is obtained by setting the price of housing $PH_{i,t}$ equal to $MC_{i,t}$:

$$I_{i,t} = H_{i,t-1} \times max\left[0, \left(\frac{PH_{i,t}}{C_{i,t}}\right)^{\varphi_{i,t}} - 1\right] \tag{A.2}$$

Assuming that the supply elasticity is always greater than zero, it follows from Eq. A.2 that investment will only take place if the price of housing exceeds the costs of construction. By inserting the expression of Eq. A.2 into Eq. A.1, and then taking logs, we get the housing supply function $S_{i,t}$:

$$S_{i,t} = \begin{cases} H_{i,t-1} & \text{if } PH_{i,t} \leq C_{i,t} \\ H_{i,t-1} + \varphi_{i,t}(PH_{i,t} - C_{i,t}) & \text{if } PH_{i,t} > C_{i,t} \end{cases} \tag{A.3}$$

The supply curve is piecewise linear and kinked: if the price of housing is smaller or equal to construction costs, the supply of homes is simply equal to the existing housing stock. If the price of housing exceeds construction costs, builders will add a flow of new construction to the existing stock. In this framework, supply is assumed to be rigid downwards, as housing is typically not demolished or dismantled (Glaeser and Gyourko 2005). Note also that supply increases linearly with the supply elasticity $\varphi_{i,t}$, as we will see below.

We specify housing demand as follows:

$$D_{i,t} = v_0 r_t + v_1 Y'_{i,t} + v_2 PH_{i,t} \tag{A.4}$$

where demand depends linearly on the interest rate r_t, assumed to be common across markets, on local house prices, and on area-specific factors captured by the vector $Y'_{i,t}$, such as household income and crime rates, as a proxy for local amenities – used before in the empirical analysis to identify a demand shift.

Consider a market where construction is greater than zero, with the equilibrium in the housing market being determined by the intersection of supply (Eq. A.3) and demand (Eq. A.4). It follows that in equilibrium, house prices and quantity of housing assume the following expressions:

$$D_{i,t} = S_{i,t}$$

$$PH_{i,t} = \frac{1}{\varphi_{i,t} - v_2}(v_0 r_t + v_1 Y'_{i,t} - H_{i,t-1} + \varphi_{i,t} C_{i,t}) \tag{A.5}$$

$$S_{i,t} = \frac{\varphi_{i,t}}{\varphi_{i,t} - v_2}(v_0 r_t + v_1 Y'_{i,t} + v_2 C_{i,t}) - \frac{v_2}{\varphi_{i,t} - v_2} H_{i,t-1} \tag{A.6}$$

We now assume that the economy is in equilibrium at time t=0, and then is hit by a positive demand shock at time t=1, say, an expansionary monetary policy shock in which the central bank reduces the interest rate r_t. The marginal impact of an expansionary monetary policy shock is given by the derivative of Eqs. A.5 and A.6 with respect to minus r_t:

$$-\frac{\partial PH_{i,t}}{\partial r_t} = -\frac{v_0}{\varphi_{i,t} - v_2} > 0 \tag{A.7}$$

$$-\frac{\partial S_{i,t}}{\partial r_t} = -\frac{\varphi_{i,t} v_0}{\varphi_{i,t} - v_2} > 0 \tag{A.8}$$

Our model predicts that both house prices and quantity would increase after an interest rate reduction, resulting from the combination of a negative numerator and positive denominator (multiplied by minus 1 as we have a reduction in the interest rate): housing demand is stimulated by declines in the interest rate (negative v_0), while supply elasticities are always equal to or greater than zero, and housing demand declines when house prices increase (negative v_2).

We illustrate the conjectures above in the left panel of Figure A.1. After a reduction in the interest rate, the demand curve shifts from D_0 to D_1, implying a new equilibrium with both higher house prices ph_1 and quantity h_1 (point B). The dotted part of the housing supply curve illustrates that housing supply is rigid downwards, so that the supply curve kinks at A at time t=0 and at B after the shock. This exercise assumes that supply elasticities are constant over time as in Green et al. (2005), Gyourko et al. (2008), Huang and Tang (2012), Glaeser et al. (2014), Anundsen and Heebøll (2016), Aastveit and Anundsen (2017). The supply elasticities only play a role over the cross-section. For instance, by exploring the variation in supply elasticities across a large sample of MSAs, Aastveit and Anundsen (2017) find that expansionary monetary policy shocks have a substantially greater impact on house prices in markets with an inelastic housing supply.

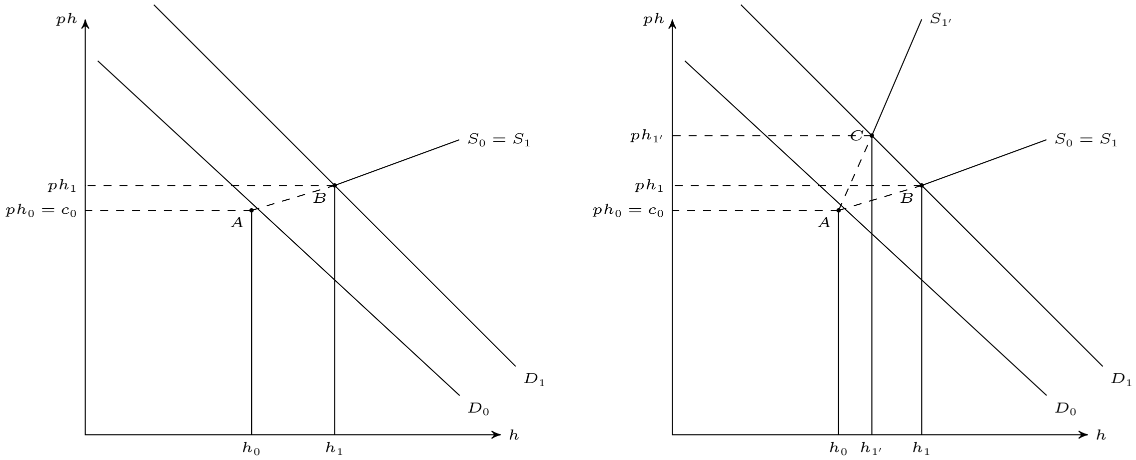
We move one step forward, and show in our model that the same logic applies within the same market: the impact of a given demand shock on house prices in the same area varies over time, if the slope of the supply curve changes between periods. When there is a reduction in the interest rate, the marginal effect of a decline in the supply elasticities on prices and quantities is given by the derivative of Eqs. A.7 and A.8 with respect to minus $\varphi_{i,t}$:

$$-\frac{\partial\left(-\frac{\partial PH_{i,t}}{\partial r_t}\right)}{\partial \varphi_{i,t}} = -\frac{v_0}{(\varphi_{i,t} - v_2)^2} > 0$$

$$-\frac{\partial\left(-\frac{\partial S_{i,t}}{\partial r_t}\right)}{\partial \varphi_{i,t}} = -\frac{v_0 v_2}{(\varphi_{i,t} - v_2)^2} < 0$$

Our model suggests that if supply elasticities decline over time for the same area, a lower interest rate would lead house prices to rise by more, and this would be reflected in a smaller expansion in supply.g We illustrate this conjecture in the right panel of Figure A.1. Assuming a decline in the supply elasticity for a given local housing market between period 0 and 1 – akin to what we have found in our empirical estimates – then a steeper supply curve implies that a demand shock moves the equilibrium to higher prices and lower quantity compared to a situation where the supply elasticity is constant (point C versus point B). In this new equilibrium, a steeper supply curve over time ($S_0 = S_1$ to $S_{1'}$) implies that a given demand shock can act as an amplification mechanism for house prices.

Figure A.1: Impact of expansionary monetary policy shock on the housing market



Market 1: Constant supply elasticity

Market 2: Steeper elasticity over time

Notes: Left panel: D_0 and S_0 are the original demand and supply curves, and point A is the initial equilibrium with house prices ph_0 and quantity h_0. After an expansionary monetary policy shock, demand shifts to D_1, and the new equilibrium is reached at point B, with both higher house prices ph_1 and quantity h_1, conditional on a time-invariant supply elasticity ($S_0 = S_1$). Right panel: If the supply elasticity declines between periods, i.e. the supply curve steepens from $S_0 = S_1$ to $S_{1'}$, the equilibrium moves to point C, with higher house prices $ph_{1'}$ and lower quantity $h_{1'}$.

Appendix B: Data description

Building permits: number of permits issued by a local jurisdiction to proceed on a construction project. Source: Census Bureau, and Moody's Analytics.

Housing starts: number of housing units in which construction work has started. The start of construction is when excavation begins for the footings or foundation of a building. Source: Census Bureau, and Moody's Analytics.

Housing stock: a house, apartment, mobile home or trailer, a group of rooms, or a single room that is occupied or available for occupancy. Updated from 2010q3 onwards by accumulating housing completions. Source: Census Bureau, and Moody's Analytics.

FHFA house price index: weighted, repeat-sales index, measuring average price changes in repeat sales or refinancings on the same single-family properties whose mortgages have been purchased or securitized by Fannie Mae or Freddie Mac. Source: FHFA, and Moody's Analytics.

UNAVAL: the land unavailability index captures housing supply geographical constraints. It is constructed using topographic maps measuring the proportion of land in a 50 km radius of the city center that is lost to steep slopes and water bodies, such as oceans, rivers, lakes and wetlands. Source: Saiz (2010).

WRLURI: the Wharton Residential Land Use Regulatory Index captures regulatory restrictions in the housing market, i.e. measures the time and financial cost of acquiring building permits and constructing a new home. It refers to the principal component of 11 survey-based measures which is interpreted as the degree of stringency of local zoning laws. Source: Gyourko et al. (2008).

Crime rates: counts of crimes per 100,000 inhabitants reported to the police for each police agency (cities, towns, and villages). It is broken down into two major types: violent crime, which includes offences of murder, forcible rape, robbery, and aggravated assault, and property crime, which includes offences of burglary, larceny-theft, and motor vehicle theft. Source: Uniform Crime Report Offenses Known to Law Enforcement dataset of the FBI.

Population: resident population in each MSA. Source: Census Bureau, and Moody's Analytics.

Population density: population per square mile. Annual data interpolated into quarterly. Data available since 2000. Source: Census Bureau, and Moody's Analytics.

CPI: consumer price index for all urban consumers. Source: BLS, and Moody's Analytics.

Disposable personal income: The income available to persons for spending or saving. It is equal to personal income less personal current taxes. Source: BEA, and Moody's Analytics.

Construction wages: wages and salaries in the construction sector. Data available at the state level. The original quarterly series has been adjusted for seasonality using X-13-ARIMA from the Census Bureau. Source: BEA.

Unemployment rate: the number of unemployed as a % of total labour force. Source: BLS, and Moody's Analytics.

Mortgage originations: dollar amount of new mortgage loans approved by the mortgage broker or loan officer. Data available until 2016q4. Source: Home Mortgage Disclosure Act, and Moody's Analytics.

Dependency ratio: ratio of people younger than 15 or older than 64 years old to the working age population (those aged 15-64). Source: Census Bureau, and Moody's Analytics.

Black: fraction of black or African American relative to total population. Annual data interpolated into quarterly. Source: Census Bureau, and Moody's Analytics.

Hispanic: fraction of people of Hispanic or Latino origin relative to total population. Annual data interpolated into quarterly. Source: Census Bureau, and Moody's Analytics.

Table B.1: Descriptive statistics

	Obs	Mean	Std. Dev.	Min	Max
Real HPI (log)	21,336	4.8	0.2	4.1	5.5
Building permits (log)	21,336	7.3	1.5	2.1	12.1
Housing starts (log)	21,336	7.3	1.4	2.2	11.6
Housing stock (log)	21,336	5.2	1.1	3.3	9.0
UNAVAL	21,336	0.3	0.2	0.0	0.9
WRLURI	21,336	-0.1	0.8	-1.8	4.3
Real personal income (log)	21,336	16.4	1.2	14.2	20.7
Real construction wages (log)	21,336	15.1	1.0	12.2	17.0
CPI (log)	21,336	5.3	0.2	4.5	5.7
Real mortgage originations (log)	19,439	13.7	1.3	8.5	18.3
Unemployment rate (%)	21,336	5.9	2.6	1.2	32.1
Population (log)	21,336	6.0	1.1	4.0	9.9
Population density	18,288	319.3	344.9	6.3	2754.3
Dependency ratio (%)	21,336	50.7	6.2	31.5	85.2
Black ratio (%)	21,272	11.7	11.2	0.1	53.9
Hispanic ratio (%)	21,272	11.3	14.7	0.4	92.2
Total crime rate (%)	17,000	3937.4	1291.2	1128.4	9469.3
Property crime rate (%)	17,360	3492.1	1159.3	3.1	8234.6
ΔReal HPI (%)	21,336	0.3	1.9	-15.7	12.3
ΔReal personal income (%)	21,336	0.6	1.3	-8.9	11.9
ΔReal construction wages (%)	21,336	0.5	3.0	-19.7	17.3
ΔCPI (%)	21,336	0.5	0.6	-3.1	4.0
ΔUnemployment rate	21,336	0.0	0.4	-8.4	6.2
ΔPopulation (%)	21,336	0.2	0.5	-44.3	10.2

Sources: Bureau of Economic Analysis, Bureau of Labor Statistics, Census Bureau, Federal Bureau of Investigation, Federal Housing Finance Agency, Gyourko et al. (2008), Home Mortgage Disclosure Act, Moody's Analytics, and Saiz (2010).

Appendix C: Details on the Bartik-type instrumental variable approach

The first stage regression when we employ the Bartik-type instrumental variable approach estimates the sensitivity of local house prices to regional house prices for each MSA and for each housing boom j:

$$\Delta \log(HPI_{i,r,t}) = \eta_i^j + \theta_i^j \Delta \log(HPI_{r,t}) + \psi_i^j X'_{i,r,t} + \epsilon_{i,r,t}^j \tag{C.1}$$

where $\Delta log(HPI_{i,r,t})$ denotes the log percentage change in house prices in MSA i of region r, and $\Delta log(HPI_{r,t})$ is the equivalent variable for the nine Census Divisions. In the spirit of the Bartik-shift share approach, our instrument for the house price variables in Eq. 6 is given by $\widehat{\theta_i^j} \Delta \log(HPI_{r,t})$. We add a set of controls $X'_{i,t}$ – construction wage growth, income growth, the change in the unemployment rate, population growth, and inflation – to minimise the potential bias arising from the possibility that local permits in our main equation may respond differentially to regional shocks through other channels than local house prices (see the discussion in Guren et al. 2018). When running the regression for MSA i, we exclude the MSA in question from the regional house price index $HPI_{r,t}$, so as to avoid biasing the coefficient θ_i^j, given that the same variable would appear simultaneously on the left and right hand sides.

After running the regression above for each MSA and for each housing boom, we collect the instrument $\widehat{\theta_i^j} \Delta \log(HPI_{r,t})$ for house prices in our supply equation Eq. 6. The coefficients are in general estimated less precisely than in the baseline regression, particularly on the interaction terms (Table D.2 in Appendix D). The Bartik-style instruments are, however, relatively weak, as the F-tests suggest that we cannot reject the null hypothesis that our instruments are not weakly correlated with the endogenous variables. One of the reasons for the low power of our instruments may be related to the difficulty of this approach in separating housing demand from supply. In addition, another reason may be related to the lack of enough time variation within each MSA to identify house prices.

Appendix D: Additionl tables and figures

Tables

Table D.1: Monetary policy shocks

	HF	Q
Mean	-0.011	-0.022
Median	0	-0.011
Std. deviation	0.067	0.067
Min	-0.413	-0.328
Max	0.125	0.128
No. Obs.	127	62

Notes: HF refers to high frequency and *Q* to quarterly.

Table D.2: Regression estimates: Bartik-type instrument

	1996-2006	**2012-2017**
$\log(HPI)$	3.895***	2.668
	(1.071)	(4.091)
$\log(HPI) \times UNAVAL$	-2.033	-6.744
	(1.807)	(5.699)
$\log(HPI) \times WRLURI$	-1.252	1.857
	(1.547)	(2.625)
$\log(H_{t-1})$	0.390***	0.214***
	(0.060)	(0.050)
Number of MSA	241	254
Observations	7,381	4,866
Cragg-Donald F-test	8.13	11.78
Kleibergen-Paap (robust) F-test	0.192	1.275

Notes: IV estimates of Eq. 6, where the dependent variable is the log of building permits. House prices have been instrumented by exploring the sensitivity of local house prices to regional house prices (see Section 7.1 for more details). The Cragg-Donald F-test and Kleibergen-Paap F-test assume that under the null the excluded instruments are not weakly correlated with the endogenous regressors. The constant and additional control variables are not reported. Robust heteroskedastic standard errors shown in parentheses. Asterisks, *, **, and ***, denote statistical significance at the 10%, 5%, and 1% levels.

Table D.3: Robustness regression estimates by housing boom

	1996-2006					2012-17				
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	Base	Tot_crime	Perm_int	SRI	Mortg	Base	Tot_crime	Perm_int	SRI	Mortg
$\log(HPI)$	2.774***	2.261***	2.671***	2.431***	2.737***	1.794**	1.730**	1.723**	1.681**	1.065
	(0.428)	(0.376)	(0.425)	(0.367)	(0.417)	(0.847)	(0.855)	(0.847)	(0.741)	(0.889)
$\log(HPI) \times UNAVAL$	-1.344***	-1.182***	-1.307***		-1.380***	-1.225	-1.359	-1.160		-2.038
	(0.340)	(0.316)	(0.336)		(0.325)	(1.185)	(1.232)	(1.184)		(1.259)
$\log(HPI) \times WRLURI$	-0.718***	-0.660***	-0.705***		-0.672***	-1.086**	-1.084**	-1.054**		-0.371
	(0.096)	(0.095)	(0.096)		(0.091)	(0.422)	(0.426)	(0.422)		(0.401)
$\log(HPI) \times SRI$				-0.432***					-0.597***	
				(0.061)					-0.193	
$\log(H_{t-1})$	0.415***	0.421***	0.424***	0.419***	0.408***	0.203***	0.201***	0.207***	0.202***	0.182***
	(0.019)	(0.018)	(0.019)	(0.019)	(0.017)	(0.023)	(0.023)	(0.023)	(0.023)	(0.021)
Number of MSA	241	241	241	241	233	254	254	254	254	242
Observations	7,548	7,464	7,548	7,548	7,442	4,866	4,758	4,866	4,866	3,812
F-test	39.83	47.73	39.99	60.53	37.81	49.66	49.33	49.71	76.53	43.87
F-test (robust)	31.00	35.12	31.19	46.74	29.75	29.61	30.09	29.65	43.20	25.77

Notes: IV estimates of Eq. 6, where the dependent variable is the log of building permits. Each column represents a separate regression: *Base* is the baseline specification, *Tot_crime* uses total crime (property crime plus violent crime) as the instrument, *Perm_int* uses permit intensity as the dependent variable, *SRI* replaces UNAVAL and WRLURI with a supply restrictions index (the sum of these two variables standardized), and *Mortg* controls for mortgage originations. The F-test and robust F-test assume that under the null the excluded instruments are not weakly correlated with the endogenous regressors. The constant and additional control variables are not reported. Robust heteroskedastic standard errors shown in parentheses. Asterisks, *, **, and ***, denote statistical significance at the 10%, 5%, and 1% levels.

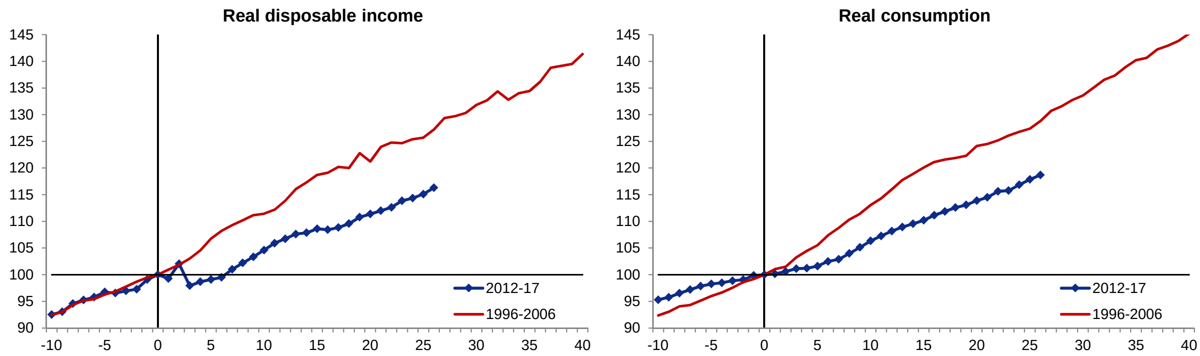
Table D.4: Estimated elasticities: alternative specifications

	1996-2006			**2012-2017**		
	p10	p50	p90	p10	p50	p90
Base	1.58	2.63	3.37	0.38	1.75	2.74
Tot_crime	1.19	2.14	2.81	0.28	1.67	2.67
Perm_int	1.50	2.53	3.25	0.36	1.68	2.64
SRI	1.48	2.51	3.26	0.36	1.79	2.83
Mortg	1.54	2.54	3.29	-0.22	0.69	1.30

Notes: Estimated elasticities from Eq. 6 for the median, 10^{th} and 90^{th} percentiles for each housing boom. *Base* is the baseline specification, *Tot_crime* uses total crime as the instrument, *Perm_int* uses permit intensity as the dependent variable, *SRI* replaces UNAVAL and WRLURI with a supply restrictions index, and *Mortg* controls for mortgage originations.

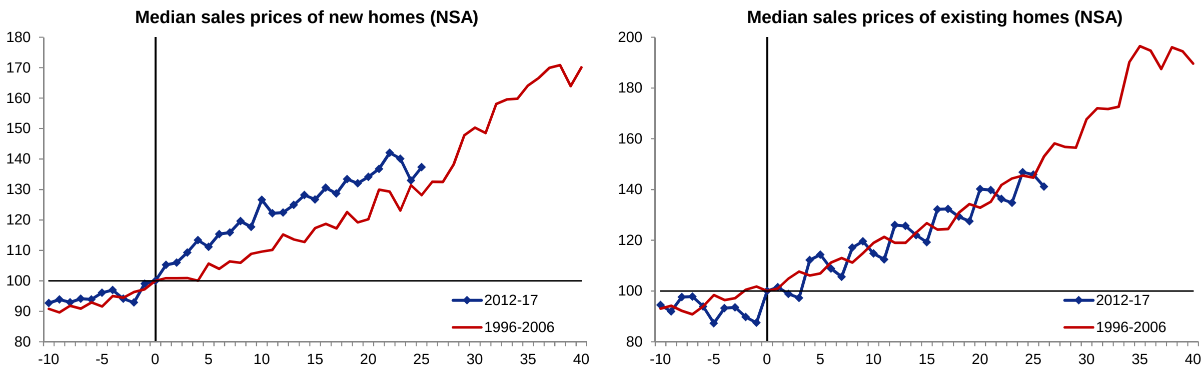
Figures

Figure D.1: Demand fundamentals across booms



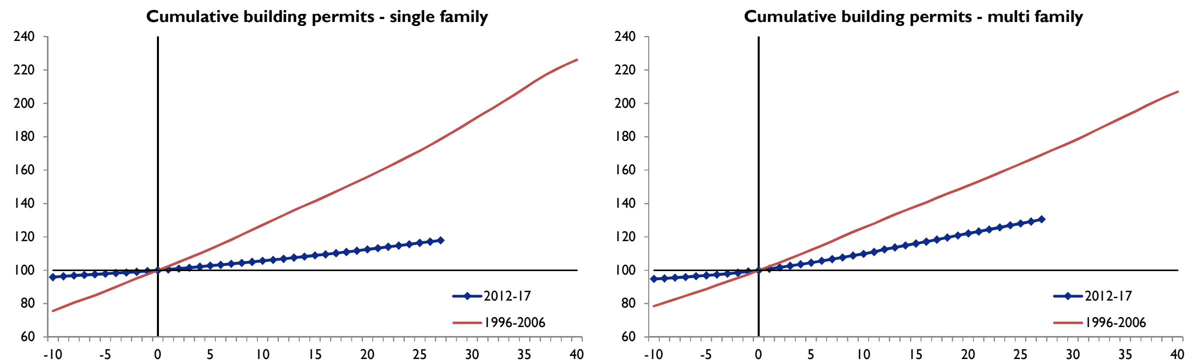
Sources: Bureau of Economic Analysis, and authors' calculations.
Notes: The figure tracks the evolution of real disposable income and real personal consumption at a quarterly frequency during the two house price booms. The zero on the x-axis marks the beginning of each housing boom. The solid line refers to the boom between 1996q4 and 2006q4, while the blue line with markers is from 2012q3 to 2017q4.

Figure D.2: Median sales prices of new and existing homes across booms



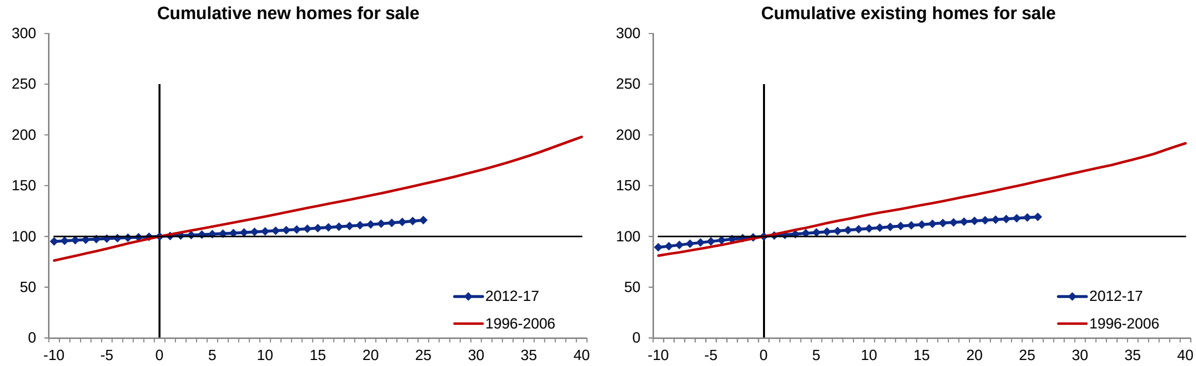
Sources: Census Bureau, National Association of Realtors, and authors' calculations.
Notes: The figure tracks the evolution of non-seasonally adjusted median sales prices of new and existing homes at a quarterly frequency during the two house price booms. The zero on the x-axis marks the beginning of each housing boom. The solid line refers to the boom between 1996q4 and 2006q4, while the blue line with markers is from 2012q3 to 2017q4.

Figure D.3: Building permits by segment across booms



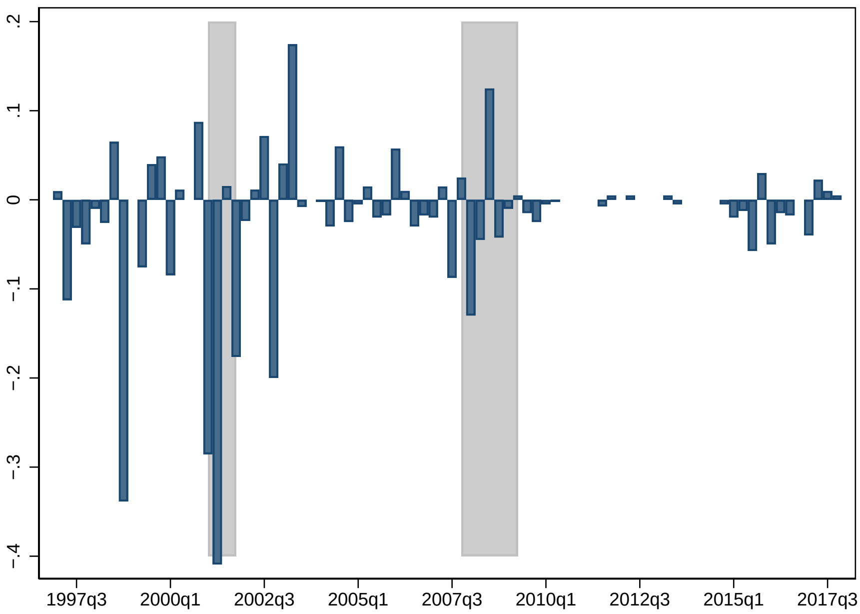
Sources: Census Bureau, and authors' calculations.
Notes: The figure tracks the evolution of single-family and multi-family building permits at a quarterly frequency during the two house price booms. The zero on the x-axis marks the beginning of each housing boom. The solid line refers to the boom between 1996q4 and 2006q4, while the blue line with markers is from 2012q3 to 2017q4.

Figure D.4: New and existing homes available for sale across booms



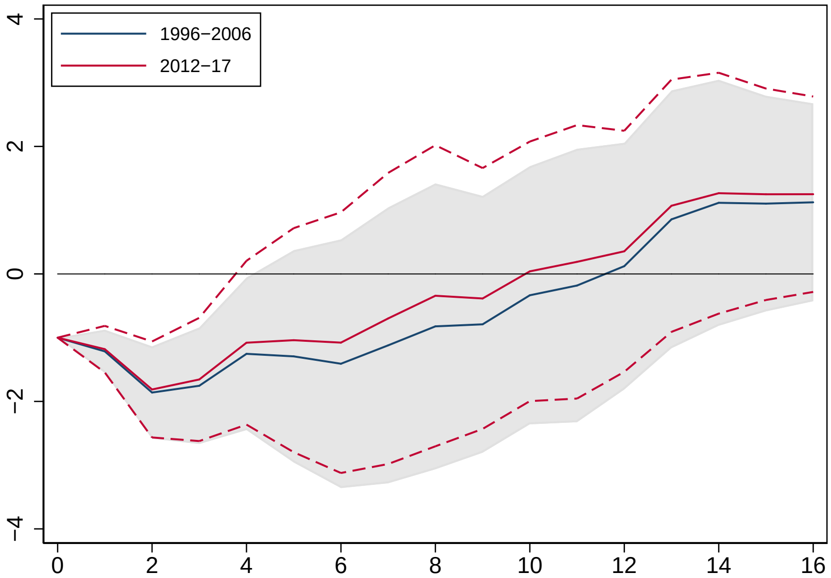
Sources: Census Bureau, National Association of Realtors, and authors' calculations.
Notes: The figure tracks the evolution of new and existing homes available for sale at a quarterly frequency during the two house price booms. The zero on the x-axis marks the beginning of each housing boom. The solid line refers to the boom between 1996q4 and 2006q4, while the blue line with markers is from 2012q3 to 2017q4.

Figure D.5: Monetary policy shocks



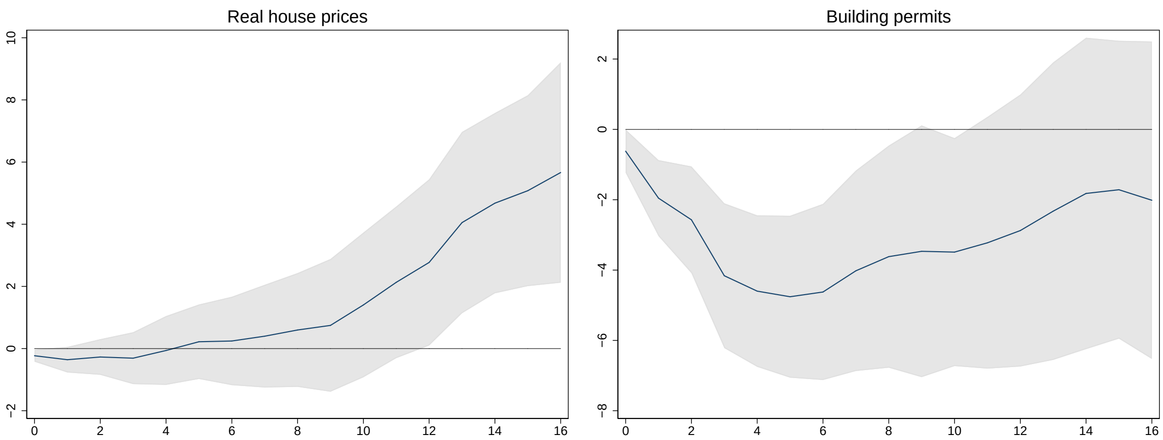
Sources: Bloomberg, and authors' calculations.
Notes: High-frequency monetary policy shocks aggregated to the quarterly frequency. Negative values refer to expansionary shocks. Shaded areas refer to recession periods, as defined by the NBER.

Figure D.6: Responses of policy indicator to an expansionary monetary policy shock



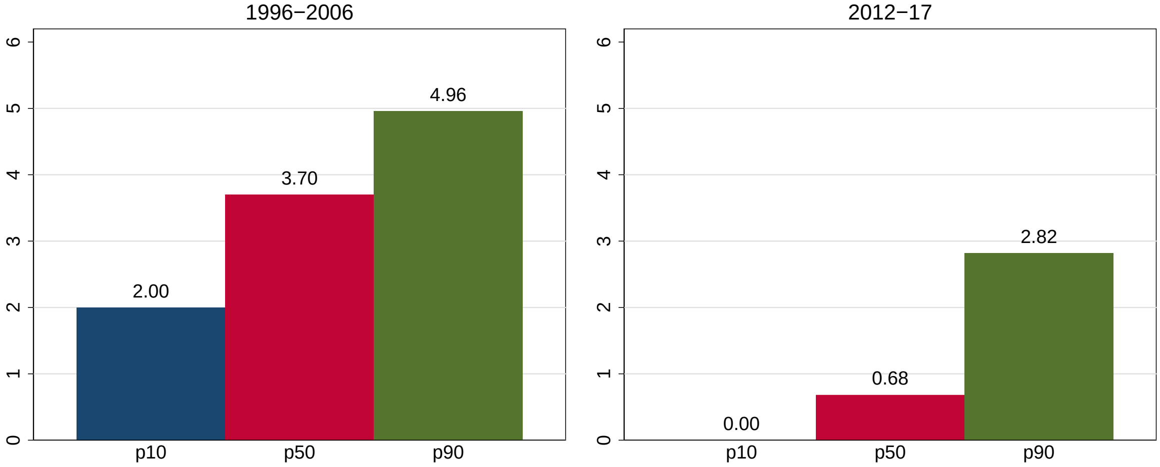
Notes: Cumulative impulse responses to a 100 basis point decline in the one-year Treasury bill yield, assessed at the sample median elasticity for each housing boom period. The grey area and the dashed red lines refer to 90 percent confidence bands.

Figure D.7: Differential effect between booms: Driscoll-Kraay estimator



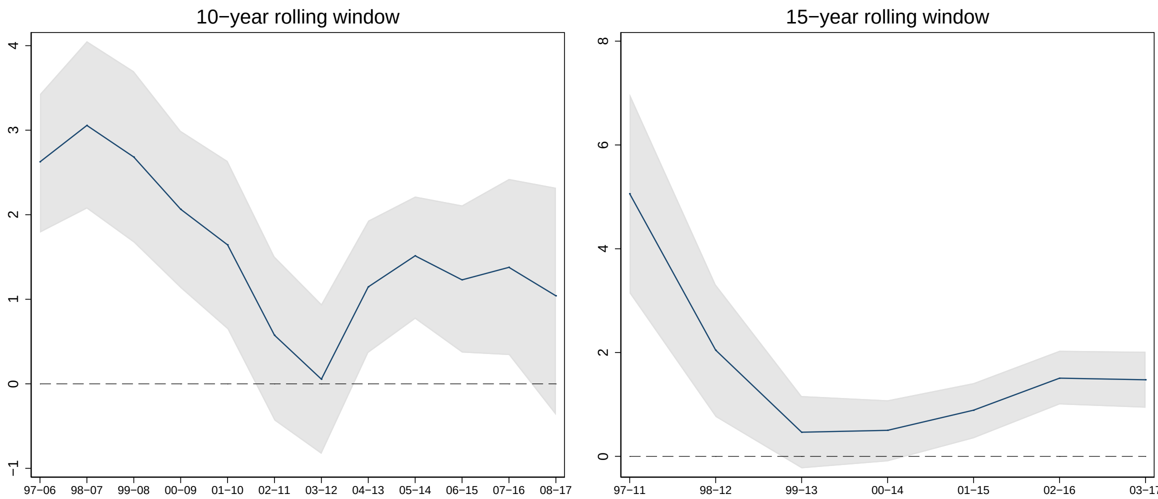
Notes: The figure depicts the difference in the estimated response of house prices and building permits between the 2012-17 and the 1996-2006 booms, with the associated 90 percent confidence bands. Standard errors have been adjusted for cross-sectional dependence in the errors across MSAs with the Driscoll-Kraay estimator.

Figure D.8: Estimated elasticities: Bartik-type instrument



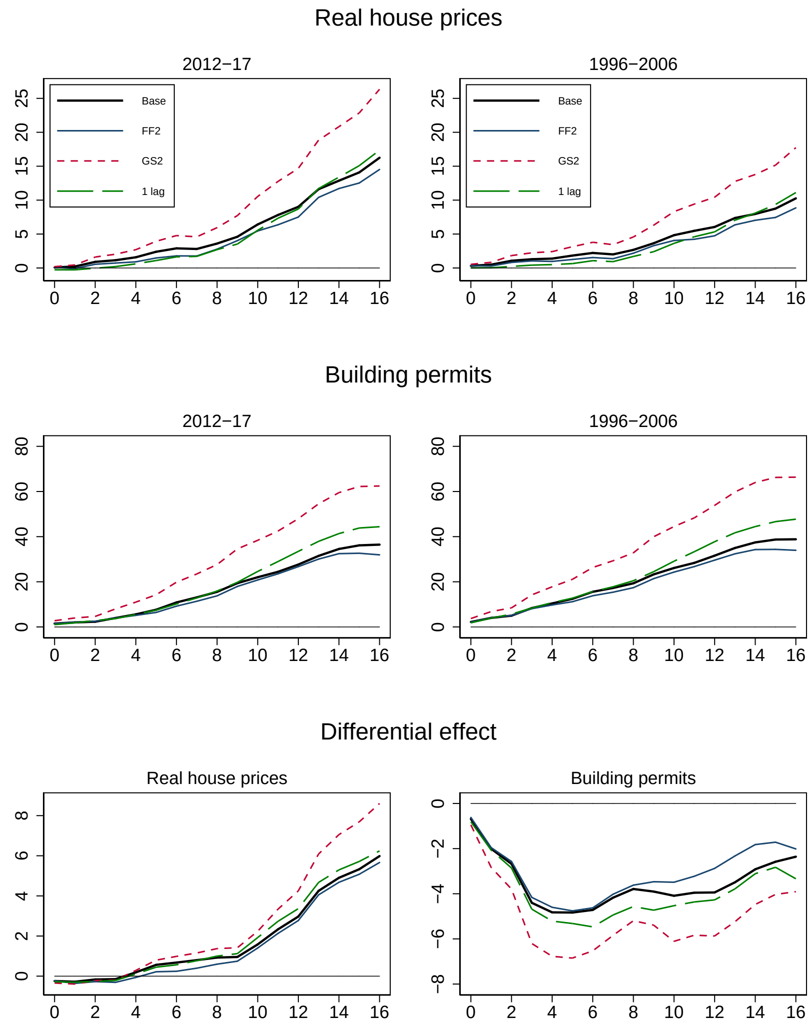
Notes: Estimated elasticities from Eq. 6 for the median, 10^{th} and 90^{th} percentiles for each housing boom. House prices have been instrumented by exploring the sensitivity of local house prices to regional house prices (see Section 7.1 for more details).

Figure D.9: Estimated elasticities with rolling windows



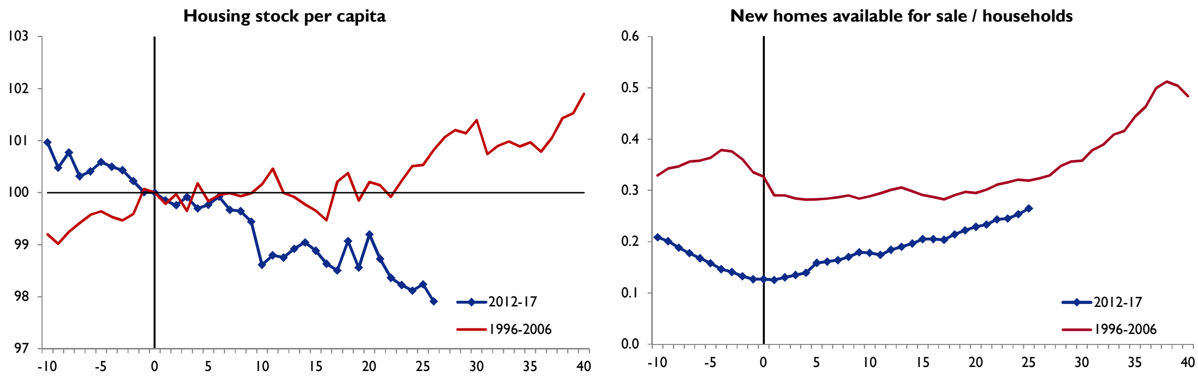
Notes: Estimated elasticities for the median and the associated 90 percent confidence bands from Eq. 6, using 10- and 15-year moving rolling windows. The x-axis refers to periods of 10 years (left figure) and 15 years (right figure).

Figure D.10: Responses to an expansionary monetary policy shock: alternative specifications



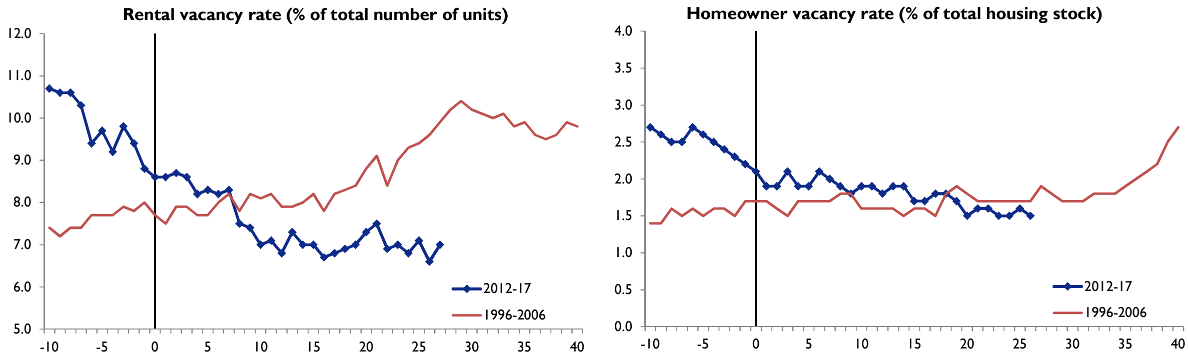
Notes: Cumulative impulse responses to a 100 basis point decline in the one-year Treasury bill yield, assessed at the sample median elasticity for each housing boom. *Base* is the baseline specification, *FF2* uses surprises in the two-month ahead Fed funds futures as the instrument for the monetary policy indicator, *GS2* uses the two-year Treasury note yield as the policy indicator, and *1 lag* is the benchmark model with only 1 lag.

Figure D.11: Housing supply indicators across booms



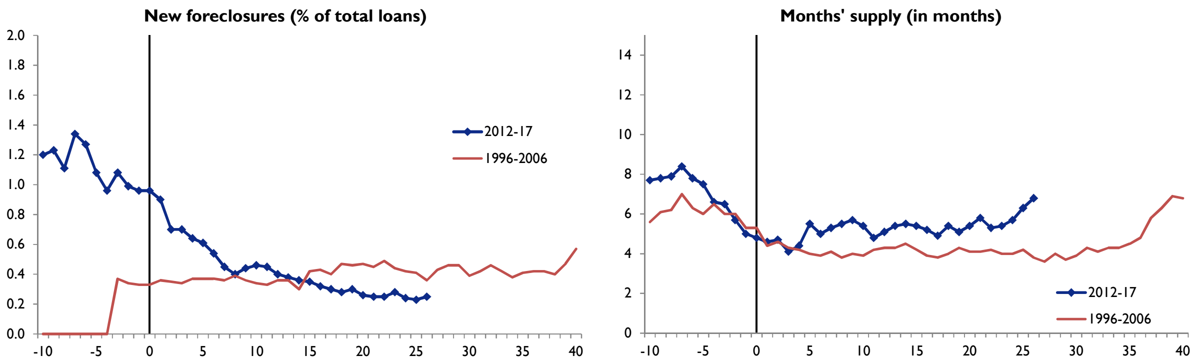
Sources: Census Bureau, and authors' calculations.
Notes: The figure shows developments in housing stock and new homes available for sale divided by the number of households during 1996q4–2006q4 (red solid line) and 2012q3–2017q4 (blue line with markers). The housing stock per capita is scaled such that it takes a value of 100 at the beginning of each period, whereas new homes available for sale per capita is displayed in level terms. The horizontal axis shows quarters around the beginning of the two booms, and the vertical line at zero is the starting point of both booms.

Figure D.12: Housing vacancy rates



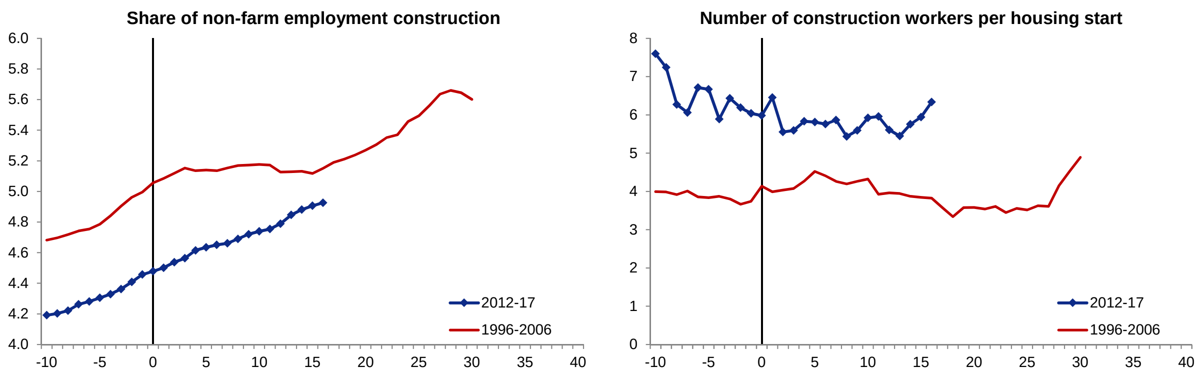
Sources: Census Bureau, and authors' calculations.
Notes: The figure tracks the evolution of housing vacancy rates at a quarterly frequency during the two house price booms. The zero on the x-axis marks the beginning of each housing boom. The solid line refers to the boom between 1996q4 and 2006q4, while the blue line with markers is from 2012q3 to 2017q4.

Figure D.13: Foreclosures and months' supply of houses



Sources: Census Bureau, CoreLogic, and authors' calculations.
Notes: The figure tracks the evolution of new foreclosures and months' supply of houses at a quarterly frequency during the two house price booms. The zero on the x-axis marks the beginning of each housing boom. The solid line refers to the boom between 1996q4 and 2006q4, while the blue line with markers is from 2012q3 to 2017q4.

Figure D.14: Construction employment across booms



Sources: Census Bureau, Bureau of Labor Statistics, and authors' calculations.
Notes: The figure tracks the evolution of the construction employment share in total employment, and the number of construction workers divided by housign starts at a quarterly frequency during the two house price booms. The zero on the x-axis marks the beginning of each housing boom. The solid line refers to the boom between 1996q4 and 2006q4, while the blue line with markers is from 2012q3 to 2017q4.