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KNUT ARE AASTVEIT
BRUNO ALBUQUERQUE
ANDRÉ K. ANUNDSEN

Changing Supply Elasticities and Regional Housing Booms

Developments in U.S. house prices over the past decade mirror those of the 1996–2006 boom. Construction activity has, however, been weak. Using data for 254 U.S. metropolitan areas, we show that housing supply elasticities have fallen markedly in recent years. We find that housing supply elasticities have declined more in areas in which land-use regulation has tightened the most, and in areas that experienced the sharpest housing busts. Consistent with the declining housing supply elasticities, we find that monetary policy shocks have had a stronger effect on house prices during the past decade than during the previous boom. At the same time, building permits respond less.

JEL codes: C23, E32, E52, R31

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Knut Are Aastveit *is a Deputy Research Director at Norges Bank and BI Norwegian Business School* (*E-mail*: knut-are.aastveit@norges-bank.no). Bruno Albuquerque *is an Economist in the Strategy, Policy, and Review Department at the International Monetary Fund and University of Coimbra* (*E-mail*: BAlbuquerque@imf.org). André Anundsen *is a Research Professor and Deputy Head of Research at Housing Lab at Oslo Metropolitan University* (*E-mail*: andre-kallak.anundsen@oslomet.no).

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At the end of 2019, real U.S. house prices had increased by around 30% since the beginning of the housing recovery in mid-2012. Recent house price developments are remarkably similar to the 1996–2006 housing boom. Despite the strong rise in house prices, construction activity has remained low and is considerably weaker than during the previous housing boom. We document that this is related to a decline in housing supply elasticities. Furthermore, we argue that there are regional differences in the extent of the decline. Against this background, we ask the following questions: (i) What factors have contributed to changing housing supply elasticities?, and (ii) How does the decline in housing supply elasticities impact the transmission of monetary policy shocks?

We consider a quarterly panel data set covering 254 U.S. Metropolitan Statistical Areas (MSAs), spanning the previous boom episode (1996–2006) and the period 2012–2019. 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 downward (Glaeser and Gyourko 2005), implying that the elasticity of housing supply should fall toward zero in a bust. For each subsample, 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 nontrivial for at least two reasons. First, there is likely reverse causality between construction activity and house prices. Second, there are large regional variations across U.S. metro areas.

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 that 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 macromodels and housing models (Dougherty and Van Order 1982, Buckley and Ermisch 1983, Meen 1990, Muellbauer and Murphy 1997, Meen 2001, 2002, Duca, Muellbauer, and Murphy 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. With respect to regional variations, 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 from Saiz (2010) and with the index of regulatory restrictions from Gyourko, Saiz, and Summers (2008).

Our main finding from the IV estimation is that U.S. housing supply elasticities have declined. We also find that there are regional differences in the extent of that decline. Housing supply elasticities may differ across areas and change over time (Green, Malpezzi, and Mayo 2005) due to changes in regulation, demographics, and in expectations about future demand and house prices. In a recent study, Herkenhoff, Ohanian, and Prescott (2018) show that there have been substantial changes in residential land-use regulation in most U.S. states over time. Looking at the change in their state-level and time-varying land-use regulation index—which is different from the one used in our IV estimation—we find that elasticities have declined the most in areas in which regulation has tightened the most. Our results also suggest a larger decline in elasticities in areas that experienced the largest drop 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.

A direct implication of lower housing supply elasticities is that a given change in housing 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, Sack, and Swanson 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 IV approach (Jordà, Schularick, and Taylor 2015, Ramey 2016, Stock and Watson 2018) to explore how monetary policy shocks affect house prices and permits in the two booms.

We find considerable heterogeneity in responses across local housing markets and over time. We estimate a substantially greater response in house prices to expansionary monetary policy shocks in supply-inelastic markets than in areas with an elastic supply. In addition, we 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 1 percentage point led to an increase in real house prices of about 5% after 3 years during the 1996–2006 boom. For the 2012–19 period, the estimated response is 10%. Consistent with this, we find that building permits increase about 15 percentage points less in response to the monetary policy shock over the same period.

Our results are robust along several dimensions. We show that the decline in housing supply elasticities is evident when: (i) employing a Bartik-type IV approach, (ii) using data on mass shootings from Pappa, Lagerborg, and Ravn (2019) and Lagerborg, Pappa, and Ravn (2020) instead of crime rates as an alternative IV, (iii) keeping the MSA fixed effects constant across the two boom periods, (iv) replacing the measures of topographical and regulatory constraints with the regulation index in Herkenhoff, Ohanian, and Prescott (2018), (v) controlling for time-varying state-specific shocks, and (vi) estimating supply elasticities using 10-year and 15-year rolling windows.

The results in this paper relate to several strands of the literature. First, a vast number of papers emphasize local differences in housing supply elasticities as a central driver of cross-sectional variation in U.S. house price developments (see, e.g., Green, Malpezzi, and Mayo 2005, Gyourko, Saiz, and Summers 2008, Saiz 2010, Huang and Tang 2012; Glaeser et al. 2014, Anundsen and Heebøll 2016).[1] This literature uses 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, Dokko et al. 2011, Williams 2011, ,2015; Jordà, Schularick, and Taylor 2015). These papers focus on the aggregate effects on house prices, which masks potential heterogeneity across regional housing markets. Notable exceptions are Fratantoni and Schuh (2003), Cooper, Luengo-Prado, and Olivei (2016) and Aastveit and Anundsen (Forthcoming), who study the effects of monetary policy on regional house prices for a samples ending in 2007 or earlier. We add to this literature by documenting nontrivial heterogeneous responses of regional house prices to a common monetary policy shock for both the 1996–2006 boom and the 2012–19 boom. Furthermore, we document a sizeable drop in housing supply elasticities over time, which have made house prices more responsive to monetary policy shocks. Paul (2020) finds that the transmission of monetary policy to house prices has become stronger since the Great Recession. Our work offers 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, Ohanian, and Prescott (2018) argue that the stronger tightening of residential land-use regulation in highly productive states, particularly California and New York, has restricted 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 U.S.

1. Van Nieuwerburgh and Weill (2010) offer a different view by arguing that cross-sectional productivity differences are the main source of variation in house prices across MSAs.

cities. In addition, Glaeser and Gyourko (2018) posit that highly regulated areas are characterized by higher house prices and lower 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 lower supply elasticities, which, in turn, amplifies the responsiveness of house prices to housing demand shocks.

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 2012–19 boom. We present the data used in our analysis, discuss our econometric approach, and estimate local housing supply elasticities for the two boom periods in Section 2. In Section 3, we explore the factors that have led to declining housing supply elasticities. Section 4 contains several robustness exercises for estimating MSA-specific housing supply elasticities. In Section 5, we analyze how changing supply elasticities affects the transmission of monetary policy shocks to the housing market. Section 6 concludes the paper.

1. THE 1996–2006 BOOM VERSUS THE 2012–2019 BOOM

To date booms and busts over the housing cycle, we analyze peaks and troughs in real house prices at the median.[2] For ease of illustration, we plot the national house price index, together with the median, the 10th and 90th percentiles of the house price distribution at the MSA level in Figure 1.

We detect three phases of the housing cycle: a strong boom from 1996 until 2006, followed by a severe bust lasting until 2012. By 2012, a new boom had started. With our data set, we cannot identify either a boom or a bust over the 1986–96 period. 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 10-year period.[3] All of the MSAs experienced increasing house prices during the 1996–2006 boom, but dispersion was high; house prices increased by 17%, on average, for the MSAs belonging to the first decile, while they increased by 97% for the top decile. During the 2006–12 bust, house prices fell in all, but one, MSA. By the end of 2019, house prices had increased in more than 98 % of the MSAs since the trough of 2012. We provide a more detailed overview of the regional dispersion in Table A.1 of the Supplementary Appendix.

2. Two alternative approaches of defining housing cycles are to identify local house price booms and busts (Ferreira and Gyourko 2011) and clustering MSAs with similar cyclical patterns (Hernández-Murillo, Owyang, and Rubio 2017).

3. In a sample of 79 MSAs, Glaeser, Gyourko, and Saiz (2008) identify a national boom over 1982–89, a subsequent bust until 1996, and a strong boom between 1996 and 2006. We get a different picture for 1986–96, since we cover a substantially larger sample of MSAs.

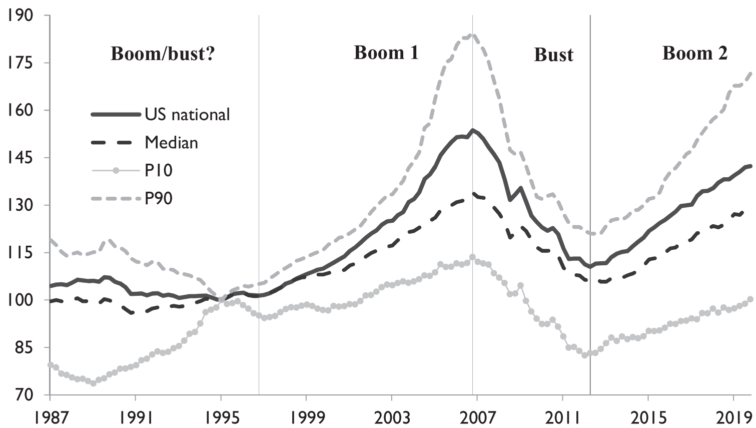

Fig 1. Real House Price Cycles.

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 U.S. aggregate index, the long-dashed blue line represents the median for the MSA distribution, the yellow line with markers represents the 10th percentile, and the dashed green line represents the 90th percentile. The vertical lines divide the sample period by phases of the housing cycle.

The dynamics of real house prices during the recent boom is similar to that of the previous housing boom. This is illustrated in the upper left panel of Figure 2, where we plot real house prices for both the 1996–2006 boom (solid line) and the 2012–19 boom (line with markers). We have scaled the house 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 2012–19 boom looks far stronger relative to income than the previous boom.[4] Although our house price index is a weighted repeat-sales index, which measures average price changes in repeat sales or refinancing on the same properties, Figure A.2 in the Supplementary Appendix documents the same pattern in house prices across booms for new homes.

Despite similar—or even stronger—developments in house prices, housing supply has grown substantially less during the 2012–19 boom.[5] During the first 7–8 years of the previous boom, housing supply—measured as the ratio of permits or starts to

4. The strong developments in house prices relative to income per capita can be partially attributed to subdued income and consumption growth, see Figure A.1 in the Supplementary Appendix for details.

5. Housing is characterized by important regional heterogeneities and variations in other attributes, see Piazzesi and Schneider (2016) and Ferreira and Gyourko (2012). The same pattern of higher house price growth and a slower supply response in the recent boom is observed in the regional data—see in particular Figure A.3 in the Supplementary Appendix.

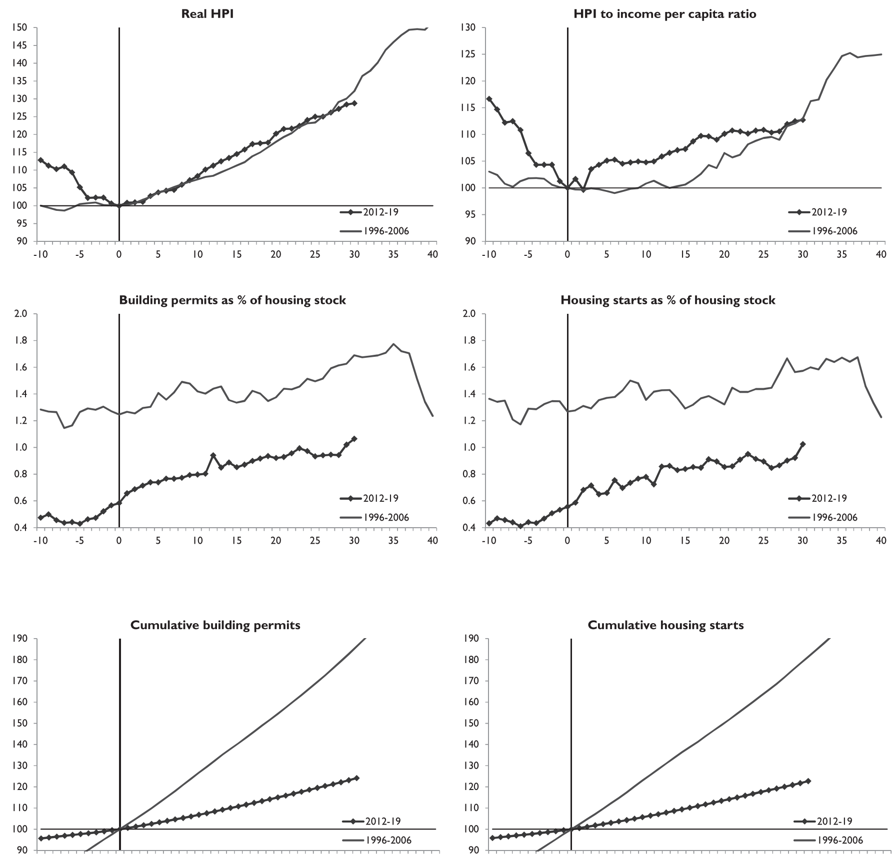

Fig 2. House Price Developments across Booms.

NOTES: The figure shows developments in real house prices, house prices relative to income per capita, building permits, and housing starts during 1996q4–2006q4 (solid line) and 2012q3–2019q4 (line with markers). The figures in the upper panel are scaled such that house prices take a value of 100 at the beginning of each period. The figures in the lower panel show the flow of building permits and housing starts scaled by the housing stock. 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.

housing stock levels—grew by an average of about 1.4%, compared to a growth of just 0.8% per year between 2012 and 2019 (lower panel of Figure 2).[6]

6. The decline in housing supply between the two booms has been driven by the single-family segment, possibly related to the shift toward urban multifamily units after the Great Recession (see Figure A.4 in the Supplementary Appendix).

2. ESTIMATING HOUSING SUPPLY ELASTICITIES IN BOOMS

2.1 Data

We use quarterly data for a panel of 254 MSAs between 1996 and 2019. The sample covers more than 80% of U.S. income and population. Our MSA definitions follow the new delineations issued by the Office of Management and Budget (OMB) from the 2010 Census. The MSA data include housing supply measures (building permits, housing starts, and housing stock), house price indices,[7] and controls for macro, financial, and sociodemographic conditions: personal disposable income, unemployment rates, mortgage originations, population, migration rates, 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 in the total population. We use wages and salaries in the construction sector at the state level to proxy builders' costs. We deflate nominal macro series with the MSA consumer price index (CPI).[8] The data have been provided by Moody's Analytics, with the original data coming from the Census Bureau, Bureau of Economic Analysis (BEA), Bureau of Labor Statistics (BLS), and Federal Housing Finance Agency (FHFA). The exception is crimes rates, which we compiled from publicly available FBI reports. The appendix contains variable definitions, and descriptive statistics for the full sample (Table A.1), the first housing boom (Table A.2), and the second boom (Table A.3).

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). Saiz (2010) uses GIS and satellite information over 1970–2000 to calculate the share of land in a 50 km 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 metro 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, Saiz, and Summers (2008). WRLURI measures the stringency of local zoning laws, that is, the time and financial cost of acquiring building permits and constructing a new home. The index is based on a nationwide survey in 2005, and on a separate study of state executive, legislative, and judicial activity. It is computed from 11 subindices measuring different types of complications and regulations when getting a building permit. It is available at a town (or city) level, which we have aggregated to the MSA level using the sample probability weights of Gyourko, Saiz, and Summers (2008).

7. Throughout the paper, the terms house prices and house price indices are used interchangeably.

8. The BLS constructs CPI data only for a selected group of MSAs. In our analysis, we use the estimates constructed by Moody's, which takes the implicit price deflators of the BEA as a starting point. Monthly data are constructed using information on cost of living, population, and median household income, and back-casting techniques.

Our sample is an unbalanced panel of 254 MSAs, since property crime rates are missing for some MSAs in certain years. In total, we have 7,546 MSA-quarter observations over the period 1997q1–2006q4, and 6,204 MSA-quarter observations over the period 2012q3–2019q4.

2.2 *Main Specification*

To estimate local housing supply elasticities across the two housing booms, we use a single-equation approach in the spirit of Green, Malpezzi, and Mayo (2005). We use the log level of building permits (sum of single-family and multifamily units) as our housing supply variable to capture the immediate reaction of builders to a change in house prices. We follow Glaeser, Gyourko, and Saiz (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. As a first exploration of how housing supply elasticities may have changed, we estimate the following specification separately for the two boom periods:

$$\log(H_{i,t}) = \rho^j \log(H_{i,t-1}) + \beta^j \log(HPI_{i,t}) + \gamma^j X'_{i,t} + \eta_i^j + \zeta_t^j + \epsilon_{i,t}^j, \tag{1}$$

in which $\log(H_{i,t})$ denotes the logarithm of building permits, and $\log(HPI_{i,t})$ the logarithm of the FHFA house price index deflated by CPI. $X'_{i,t}$ is a vector of local economic, financial, and sociodemographic variables. The economic and financial controls include the logarithm of real construction wages, which acts as a proxy for construction costs, the unemployment rate, the inflation rate, and the change in the logarithm of mortgage originations (the amount of new mortgage loans) to account for subdued credit developments since the Great Recession. To account for sociodemographic changes, we include the log of population, net domestic and international migration rates, the dependency ratio, and the fraction of Blacks and Hispanics in the total population. To control for the possibility that builders may have a preference for constructing housing in areas that have a more stable income base, we include the 5-year rolling window of the coefficient of variation in per capita income and lagged per capita income growth as additional controls. We also include the lagged dependent variable $\log(H_{i,t-1})$ to control for persistence effects. 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, $j = \{1996 - 2006, 2012 - 2019\}$. The parameters of interest are the β^j's, which tells us whether there is a change in average housing supply elasticity across the two subsamples.

Having estimated average supply elasticities for the two boom periods, we explore regional differences in supply elasticities both within and across periods. For this purpose, we estimate the following specification separately for the two boom periods:

$$\begin{aligned}\log(H_{i,t}) =& \rho^j \log(H_{i,t-1}) + \beta^j \log(HPI_{i,t}) + \lambda^j\big[\log(HPI_{i,t}) \times UNAVAL_i\big] \\ &+ \delta^j\big[\log(HPI_{i,t}) \times WRLURI_i\big] + \gamma^j X'_{i,t} + \eta_i^j + \zeta_t^j + \epsilon_{i,t}^j,\end{aligned} \tag{2}$$

in which $UNAVAL_i$ the land unavailability index of Saiz (2010) and $WRLURI_i$ the Wharton Land Use Regulatory Index (Gyourko, Saiz, and Summers 2008).

The interaction terms in equation (2) implies that housing supply elasticities may differ across MSAs if there are differences in land availability or regulatory barriers. 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 a given MSA in housing boom j is found by differentiating equation (2) with respect to house prices:

$$Elasticity_i^j = \beta^j + \lambda^j \times UNAVAL_i + \delta^j \times WRLURI_i. \tag{3}$$

2.3 IV Identification

To deal with reverse causality between house prices and permits, we use an IV approach. We use two instruments for house prices that lead to shifts in housing demand (relevance), but that does not shift housing supply (exogeneity). The first instrument exploits variation in crime rates across MSAs and over time.[9]

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). Crime rates should then capture exogenous variation in (negative) amenities that drive house price changes both within and across MSAs.

Since crime rates in the United States have been declining over time, one may question whether there is sufficient variation in crime rates over time and across MSAs to justify their use in our framework. To the extent that the decline is similar across MSAs, the rank order would be unchanged, and crime rates would almost act as a fixed effect. There are, however, large differences in crime rates over time and across MSAs in our sample. When sorting MSAs by property crime rates over time, we find substantial changes in the rank order of the MSAs between the two booms, with a standard deviation of 36, ranging from a minimum of −71 to a maximum of 118. For example, a quarter of the MSAs have experienced an increase in property crime rates between the two booms *relative* to at least 40 other MSAs. For more details, see Tables B.1 and B.2 in the Supplementary Appendix.

The relevance condition is supported by findings in the literature that point to high crime rates being strongly and negatively associated with property prices (Thaler

9. A similar instrument is used by Dell'Ariccia, Igan, and Laeven (2012) to study the link between lending standards and credit booms around the Great Recession.

1978, Schwartz, Susin, and Voicu 2003, Gibbons 2004, Pope and Pope 2012). We use property crime, which accounts for almost 90% of total crime, as our main measure of crime since it is available for a larger sample of MSAs compared with violent crime.

Our second instrument 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, Muellbauer, and Murphy 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, that is, leading to movements along, but not shifts in, the supply curve. One concern is that housing supply conditions may be endogenous to property crime, invalidating the use of our instrument. More affordable housing may be associated with higher (property) crime, implying a negative association between crime and house prices. But high-income neighborhoods may be more prone to property crime, implying a positive association. 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). One potential concern is that differences in crime rates (or income developments) may affect productivity in the construction sector. This would make the instrument(s) correlated with the residuals in the supply equation, and therefore, violate the exclusion restriction. To shed some light on this, we look at the association between our instruments and wage growth in the construction sector. Table 1 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, or for the full sample. In the particular case of crime rates, our results suggest that crime is expected to be mostly capitalized in land values, since labor is mobile, and not to show up on the supply side of the economy (wages and housing supply).

One may also argue that developments in current income contain relevant information for forecasting future demand, which could drive construction activity. For instance, income in a specific area could predict how the economy of that area may evolve over time relative to current conditions. This would violate the exclusion restriction that personal income only affects housing supply through house prices. To proxy for future demand, we consider the change in the unemployment rate h periods ahead relative to the current unemployment rate and the h periods ahead change in

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.

TABLE 1

Relationship between Wage Growth and House Price Instruments

	1996–2006	2012–2019	1996–2019
log(Crime)	−0.02	0.01	−0.04
	(0.12)	(0.13)	(0.07)
log(Income)	−0.02	−0.02	−0.00
	(0.02)	(0.02)	(0.01)
Observations	7,546	6,204	18,956
Number of MSAs	241	254	254
Adj. R^2	−0.01	−0.01	−0.01

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

TABLE 2

Relationship between Future Demand and Current Income

Panel A: Δ*Unemployment Rate*:	$h = 4$	$h = 8$	$h = 12$	$h = 16$
log(Inc.)	0.27	0.11	−0.35	−0.03
	(0.30)	(0.47)	(0.56)	(0.60)
Observations	23,155	22,211	21,267	20,251
Number of MSAs	254	254	254	254
Adj. R^2	0.29	0.38	0.51	0.58
Panel B: Δ*Employment to Population Ratio*:	$h = 4$	$h = 8$	$h = 12$	$h = 16$
log(Inc.)	−0.00	−0.01	−0.02	−0.03*
	(0.01)	(0.01)	(0.01)	(0.01)
Observations	22,211	21,267	20,251	19,235
Number of MSAs	254	254	254	254
Adj. R^2	0.14	0.21	0.30	0.40
Panel C: Δ*Housing Vac. Rates*:	$h = 4$	$h = 8$	$h = 12$	$h = 16$
log(Inc.)	0.65	1.43	1.97*	1.47
	(0.76)	(0.95)	(1.02)	(1.06)
Observations	4,504	4,148	3,780	3,412
Number of MSAs	71	71	71	64
Adj. R^2	0.41	0.45	0.49	0.53

Note: OLS estimates over 1997q1–2019q4, with MSA-fixed effects and time effects. In Panel A, we show results when the dependent variable is the change in the unemployment rate for horizons 4 (column 1) to 16 (column 4) relative to the contemporaneous unemployment rate. Panel B and Panel C show results when the dependent variable is the *h* period ahead change in employment-to-population ratio and the housing vacancy rate, respectively. 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.

employment-to-population ratios. Panels A and B of Table 2 show that we do not find any statistically significant association between income and future demand, as measured by the correlation between log of income and the different leads of either the change in the unemployment rate or the change in employment-to-population ratios, respectively. One exception is that for h=16, our results suggest a negative association between income and the change in employment-to-population ratios at the 10% level of significance. In Panel C of Table 2, we use housing vacancy rates as an alternative proxy for future demand. Data for vacancy rates are only available since 2005q1, and only cover 71 MSAs. With this caveat in mind, our results show that income does not contain any predictive power for future developments in housing vacancy rates, with the exception of h=12, for which our results suggest a positive association at the 10% level of significance. This result generally reinforces our assumption that the exclusion restriction is valid.

2.4 IV Specification

In the uninteracted model, we have one endogenous regressor (house prices) and two instruments (crime rates and income). For each boom, we estimate the following first- and second-stage regressions:

$$HPI_{i,t} = \rho_1^j \log(Crime_{i,t}) + \rho_2^j \log(Inc_{i,t}) + \phi^j X'_{i,t} + \psi_i^j + \nu_t^j + \mu_{i,t}^j, \tag{4}$$

$$\log(H_{i,t}) = \rho^{IV,j} \log(H_{i,t-1}) + \beta^{IV,j} \log(\widehat{HPI_{i,t}}) + \gamma^j X'_{i,t} + \eta_i^j + \zeta_t^j + \epsilon_{i,t}^j, \tag{5}$$

in which j signifies that all parameters may differ between the two booms. To control for possible confounders, we add a set of control variables, listed in Section 2.2. We assess the relevance and strength of the instruments with the Cragg–Donald F-statistic test for weak identification, 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% relative bias to test for weak instruments.

Results are reported in the first two columns of Table 3 for both the 1996–2006 boom and the 2012–19 boom. We use Conley (1999, 2008) standard errors that are robust to both spatial correlation and autocorrelation, by employing the code developed by Hsiang (2010). We use the QGIS-software to calculate latitudes and longitudes of MSA centroids, and set the cutoff distance for the spatial correlation at 100 miles. The kernel that is used to weigh the spatial correlations decays linearly with distance in all directions. Results show a substantial decline in the supply elasticity. Specifically, we find that the average supply elasticity of 1.63 in the first boom declines to 0.80 in the second boom. We therefore find evidence that the average housing supply elasticity has declined considerably.

To explore regional differences in supply elasticities within and across boom periods, we employ a similar IV approach. In this case, we will have three endogenous regressors: $\{HPI_{i,t}, HPI_{i,t} \times UNAVAL, HPI_{i,t} \times WRLURI\}$, since the interacted variables are also endogenous. We therefore have *six* instruments; the uninteracted instruments, as well the interactions of crime rates and income with the two supply restriction indices.

Results from the interacted model are displayed in columns 3–4 of Table 3. We find that the coefficient on house prices is statistically significant at conventional lev-

TABLE 3

Regression Estimates by Housing Boom

	No Interactions		Interactions	
	1996–2006	2012–2019	1996–2006	2012–2019
log(HPI)	1.63***	0.80*	3.35***	0.97*
	(0.28)	(0.42)	(0.49)	(0.58)
log(HPI) × UNAVAL			−1.88***	−0.09
			(0.36)	(0.95)
log(HPI) × WRLURI			−0.73***	−0.55**
			(0.09)	(0.26)
log(H_{t-1})	0.47***	0.27***	0.40***	0.27***
	(0.02)	(0.02)	(0.02)	(0.02)
Observations	7,546	6,204	7,546	6,204
MSAs	241	254	241	254
Cragg–Donald *F*-test	158.67	316.16	30.21	77.33
Kleibergen–Paap (robust) *F*-test	73.84	168.37	21.13	48.93

Note: IV estimates of equation (5), where the dependent variable is the log of building permits. The first two columns report results based on a specification with no interaction effects. The next two columns show results when we add interaction effects with WRLURI and UNAVAL. 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. We use Conley (1999, 2008) standard errors that are robust to both spatial correlation and autocorrelation, by employing the code developed by Hsiang (2010). We use the QGIS-software to calculate latitudes and longitudes of MSA centroids, and set the cutoff distance for the spatial correlation at 100 miles. The kernel that is used to weigh the spatial correlations decays linearly with distance in all directions. The standard errors are reported in absolute value in parenthesis below the point estimates. Asterisks, *, **, and *** denote statistical significance at the 10%, 5%, and 1% levels.

els, and positive, for both housing booms. Similar to the model without interaction effects, there is evidence of a statistically significant 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. We have tested for equal supply elasticities across the two booms when evaluating at the 10th, 50th, and 90th percentiles of the WRLURI and UNAVAL distributions. We reject the hypothesis of equal supply elasticities in all cases (*p*-values are 0.004, 0.003, and 0.039, respectively), implying that there is a statistically significant decline in estimated elasticities across the two booms. The interaction of house prices with the supply restriction variables yields the expected negative signs, that is, 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 last part of the sample.

In both the uninteracted and in the interacted model, the first-stage *F*-test and robust *F*-test are well above Stock and Yogo (2005)'s threshold value, suggesting that our instruments are valid and strong.

2.5 *Estimated Elasticities*

We calculate MSA-specific elasticities for the two booms by inserting the relevant parameters of equation (5) into the expression of equation (3). This gives us an estimated elasticity for each MSA for both housing booms. Figure 3 shows the estimated elasticities. The horizontal lines refer to the median, the whiskers show the 10th–90th

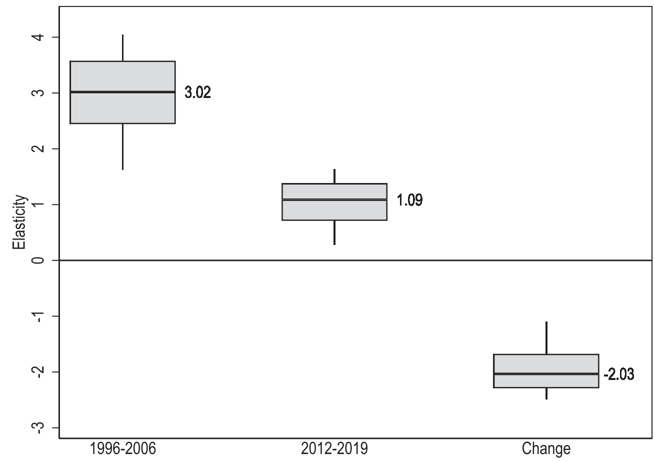

Fig 3. Estimated Elasticities: IV Specification.

NOTES: The first two bars show the distribution of estimated elasticities for each housing boom, while the last bar shows the change in the elasticities between 2012–19 and 1996–2006. The horizontal lines refer to the median, the whiskers show the 10th–90th percentiles, while the shaded blue area covers the interval between the 25th and 75th percentiles.

percentiles, while the shaded blue area covers the interval between the 25th and 75th percentiles. Our results suggest that supply elasticities have fallen across the whole distribution.[11]

In Figure 4, we show that the largest percentage decline in elasticities between the two booms, given by the darker colors, generally has taken place in the areas with the lowest elasticities during the first housing boom.

3. 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, Ohanian, and Prescott (2018) documents a substantial tightening in land-use policy in most U.S. states since 1950. They find that a substantial tightening across states took place between 1990 and 2014, of around 18%. 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, Glaeser and Gyourko 2018).

11. Maps of estimated elasticities for each MSA for the two booms are shown in Figure B.1 in the Supplementary Appendix.

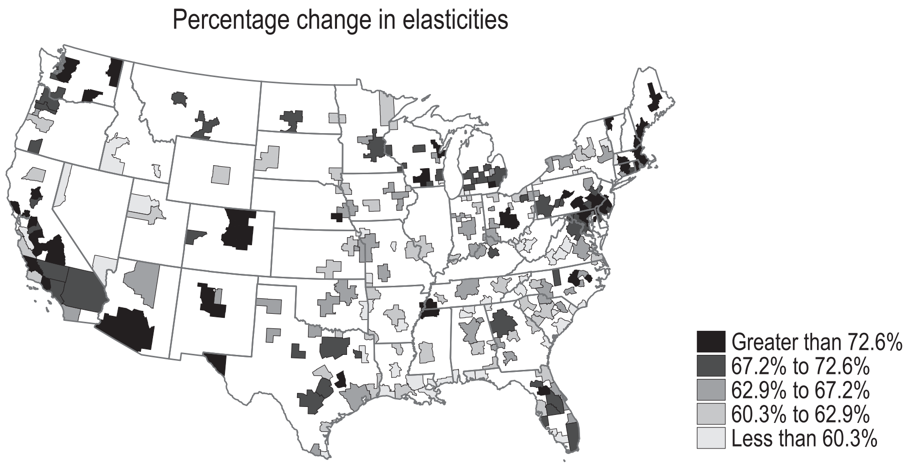

Fig 4. Percentage Change in Estimated Elasticities between Booms.

NOTES: The darker (lighter) colors refer to the largest (smallest) percentage declines in the estimated elasticities between 2012–19 and 1996–2006.

A simple correlation analysis between our estimated elasticities and Herkenhoff, Ohanian, and Prescott (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).[12] We show that this relationship holds in a multivariate setting, by estimating the following cross-sectional equation:

$$\frac{(\widehat{Elast}_i^{19} - \widehat{Elast}_i^{06})}{\widehat{Elast}_i^{06}} = \alpha + \beta_1 \frac{(X_i^{19} - X_i^{06})}{X_i^{06}} + \beta_2 Z_i + \epsilon_i, \tag{6}$$

where the dependent variable is the percentage change in estimated elasticities between 2012–19 ($\widehat{Elast}_i^{19}$) and 1996–2006 ($\widehat{Elast}_i^{06}$), as computed from equation (3) for each MSA. We regress this variable on the percentage change in a set of indicators X_i for the same period. These indicators consist of the state-level land-use regulation index of Herkenhoff, Ohanian, and Prescott (2018), in which a lower value represents a tightening of regulation, population density, construction wages, unemployment rate, and the change in the inflation rate. Finally, we control for initial conditions, Z_i, including population density and house prices to income per capita. We compute the ratio of house prices relative to income per capita by dividing the index of house

12. Herkenhoff, Ohanian, and Prescott (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 2012–19 booms.

TABLE 4

WHAT FACTORS ARE ASSOCIATED WITH THE EXTENT OF THE DECLINE IN ELASTICITIES BETWEEN BOOMS?

	(I)	(II)	(III)
Δ Land reg.	0.05**	0.05**	0.05**
	(0.02)	(0.02)	(0.02)
Δ HPI^{06-12}		0.16***	0.09*
		(0.04)	(0.05)
Δ Pop. den.			−0.07
			(0.06)
Δ Wage			0.04
			(0.03)
Δ Unemp.			1.25
			(0.85)
HP-to-inc.^{2012}			0.03
			(0.07)
Pop. den.^{2012}			−0.01**
			(0.00)
Observations	252	252	252

NOTE: Regression estimates of equation (6), where the dependent variable is the percentage change in the estimated supply elasticities between 2012–19 and 1996–2006. Robust heteroskedastic standard errors in parentheses produced by bootstrapping 1,000 replications. Asterisks, *, **, and *** denote statistical significance at the 10%, 5%, and 1% levels.

prices by an index of personal income per capita. Both the numerator and the denominator have been rebased to the same base period (100=1995q1). We also include the cumulative change in house price growth during the 2006–12 bust. To account for the generated regressor issue from the estimated elasticities, $\widehat{Elast}_i$, we bootstrap standard errors based on 1,000 replications.

Our results suggest that tighter land-use regulation is associated with a decline in elasticities between the two booms (Table 4). We find that areas with higher initial levels of population density tend to be associated with larger declines in elasticities. In contrast, we do not find any statistical link between stronger economic performance (measured by the change in the unemployment rate) or higher initial level of house prices relative to income and lower supply elasticities. Finally, we find that the areas that experienced the strongest bust in house prices over the period 2006–12 (ΔHPI^{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 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.

4. ROBUSTNESS CHECKS AND OTHER POSSIBLE EXPLANATIONS FOR DECLINING SUPPLY ELASTICITIES

4.1 Mass Shootings as an Alternative IV

While we use crime rates as an instrument for house prices, other candidate instruments include school shootings and mass shootings. For instance, Pappa, Lagerborg, and Ravn (2019) show that school shootings, and mass shootings with at least four fatalities, are exogenous to the unemployment rate. In addition, Lagerborg, Pappa, and Ravn (2020) use mass shootings with at least seven fatalities as an instrument for consumer confidence in order to estimate the effect of sentiment shocks.

Since we are considering a panel of MSAs, we mapped the location of the different mass shooting events to our MSAs. In addition, we used the Mother Jones data set to extend the sample through 2019. We then replaced property crime rates with a dummy variable that takes the value 1 for MSA-year-quarter combinations in which a mass shooting took place, and zero otherwise. We follow Pappa, Lagerborg, and Ravn (2019) and Lagerborg, Pappa, and Ravn (2020) and exclude mass shootings with personal relations. We end up with 52 events that can be mapped to the MSAs included in our data set over the sample period we consider. These events are spread throughout 39 of our 254 MSAs. Our results are robust to this alternative instrument. Detailed results showing the distribution of estimated elasticities are shown in Figure C.1 in the Supplementary Appendix.

4.2 Constant MSA Fixed Effects across Subperiods

Our baseline approach is to estimate a split-sample model to allow all coefficients, including the MSA fixed effects, to vary across the two boom period. One could argue that a more direct approach to testing whether housing supply elasticities have declined would be to hold the estimated fixed effects constant across periods and only allowing the parameters of interest to vary. We have explored the robustness of our results to estimating a specification in which the MSA fixed effects are constant across the two boom periods. Our main finding that housing supply elasticities fall in the second housing boom is maintained. Detailed results showing the distribution of estimated elasticities are shown in Figure C.2 in the Supplementary Appendix.

In our baseline specification, we allow MSA fixed effects to vary across the two boom periods for three reasons. First, the Global Financial Crisis (GFC) may have had an heterogeneous impact across MSAs. For instance, Mian and Sufi (2014) have shown that U.S. counties that suffered greater declines in housing net worth also experienced larger falls in nontradable employment and Mian and Sufi (2014) show that these losses in the nontradable sector may be long-lasting, due to the presence of wage and search frictions in the labor market. The fixed effects or the specific characteristics of several MSAs—particularly those areas that experienced large boom busts in the housing market—may therefore have changed since the GFC. Second, by allowing the MSA fixed effects to vary across periods, we partly capture changes in regulation over time that may affect building activity directly (on top of the change in the supply elasticity). Different MSA fixed effects across booms capture the level effect of changes in regulation. Third, studies such as Cheng, Liao, and Schorfheide (2016) and Aastveit et al. (2017) find evidence of instabilities in empirical macroeconomic models (factor models and vector autoregressive ~~[VARs]~~) around and after the GFC. If these instabilities are also present at the MSA level, a split-sample specification would be favorable.

4.3 Bartik-Type Instrumental Variable Approach

We check the robustness of our baseline estimates of housing supply elasticities to employing a Bartik-type IV approach (Bartik 1991). More specifically, we follow a similar approach as Guren et al. (2021), and instrument MSA-level house prices with house prices at the Census Division level. 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 2012–19 boom. A detailed description of the approach is provided in the Supplementary Appendix, and Figure C.3 shows the distribution of estimated elasticities.

4.4 Using the Regulation Index of Herkenhoff, Ohanian, and Prescott (2018) When Calculating Elasticities

The main advantage of UNAVAL and WRLURI is that they are available at the MSA level. In this way, we are able to compute supply elasticities for each MSA and for both housing booms. A disadvantage of instead using the regulation index of Herkenhoff, Ohanian, and Prescott (2018) is that this series is confined to the state level. In practice, we would be assigning the same change in regulation to the MSAs within the same state, even if there is considerable heterogeneity across MSAs belonging to the same U.S. state.

Nevertheless, we have tested the robustness of our results to replacing UNAVAL and WRLURI with the Herkenhoff, Ohanian, and Prescott (2018) index. Results remain qualitatively similar to our baseline estimates: there is a substantial decline in the house price coefficient over time. In addition, the coefficient on land-use regulation index is positive and statistically significant in the first boom (a decline in this regulation index represents a tightening). The resulting median supply elasticity for the 1996–2006 is estimated at 1.62, while for the second boom, it drops to 0.84, in line with our main results. Detailed results showing the distribution of estimated elasticities are shown in Figure C.4 in the Supplementary Appendix.

4.5 Controlling for State-by-Time Fixed Effects

In our baseline specification, we partially capture productivity changes in the construction sector by controlling for state-level construction wages. To allow for a more flexible approach, in which we control for all possible sources of time-varying state-level shocks, we add state-by-time fixed effects. Including state-by-time fixed effects in the specification introduces the challenge of having enough MSA-level variation within each state to estimate MSA-specific housing supply elasticities. For states with few MSAs, the state-by-time fixed effect will either fully—for states in which our sample only covers one MSA—or partly capture MSA-specific house price movements. This will contaminate the estimates of housing supply elasticities. Therefore, to ensure enough MSA variation within each state over time, it would be preferable to exclude states with few MSAs. This comes at a cost, however, since it reduces the degrees of freedom (about 30% of the MSAs in our sample belong to a state with 5 or fewer MSAs, and as much as 70% of the MSAs belong to a state with less than 10 MSAs). To explore this trade-off in some detail, we have added state-by-time fixed effects when excluding MSAs belonging to a state with less than i MSAs, where we let i run from 1 to 10. The implied drop in elasticities between the two booms is shown in Figure C.5 in the Supplementary Appendix. Our main result that elasticities have declined across the two booms is maintained in all cases.

4.6 *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. We have explored the robustness of our results to using 10- and 15-year rolling windows when estimating housing supply elasticities. For the 10-year window, the first regression covers the period 1997–2006, the second regression spans the period 1998–2007, and so on. Similarly, for the 15-year rolling windows, the first period goes from 1997 to 2011, and the last from 2005 to 2019. The rolling window estimation corroborates the finding that housing supply elasticities have declined over time. Further, the durability of housing entails that housing supply is rigid downwards (Glaeser and Gyourko 2005), implying that housing supply elasticities should fall toward zero during severe busts. Consistent with this, we find a particularly strong decline in housing supply elasticities during the recent housing bust. Full details are displayed in Figure C.6 in the Supplementary Appendix.

4.7 *Additional Robustness Checks*

We have carried out additional robustness checks to estimating housing supply elasticities: (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) using the Arellano–Bond estimator to account for the Nickell (1981) bias in dynamic panels (iv) replacing UNAVAL and WRLURI with a summary measure of supply restrictions, (v) using housing starts and the housing stock as the dependent variables; (vi) estimating the supply equation separately for multifamily building permits; (vii) clustering the standard errors at the state level, and (viii) using 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, Gottlieb, and Gyourko (2012). Our results are robust to these alternative specifications.

4.8 *Other Possible Explanations for Declining Supply Elasticities*

The sluggish recovery in housing supply could be linked to the strong rise in construction activity during the 1996–2006 boom that led to an oversupply of houses in the subsequent period (Nathanson and Zwick 2018, Rognlie, Shleifer, and Simsek 2018, Gao, Sockin, and Xiong 2020). There are, however, several reasons why we do not think this is a major explanation for the slow supply response: (i) the months' supply of houses is only slightly above the levels recorded during the previous boom (Figure C.7 in Supplementary Appendix),[13] (ii) the number of homes available for sale per capita is low (Figure C.8 in Supplementary Appendix), (iii) housing vacancy rates have shown similar developments across the two booms (Figure C.9 in Supplementary Appendix), and rental vacancy rates are below precrisis levels, and (iv) new foreclosures and delinquency rates are similar across the two booms (Figure C.10 in Supplementary Appendix).

Second, the weak response of builders during the recent housing boom could also be explained by difficulties in expanding capacity given the shortage of workers in the construction sector (JCHS 2018). Low unemployment and high job vacancy rates in the construction sector appear to corroborate this story (Bowman 2020). Similarly, the share of workers employed in the construction sector remains slightly lower than before the crisis. However, employment in the construction sector is actually above precrisis levels and the number of construction workers per housing start has increased (Leamer 2015) (Figure C.11 in Supplementary Appendix). There is some evidence that firms' difficulties in expanding capacity is an additional explanation for the decline in supply elasticities.

Third, following the implementation of Basel III under the Dodd–Frank Act, U.S. 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% to about 10%–11% for C&I loans more generally, it raised required capital to 15% 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.

Finally, the increased market concentration in the home building sector (Haughwout et al. 2013, Cosman and Quintero 2019, Davis and Haltiwanger 2019) is consistent with our finding of a nationwide decline in housing supply elasticities.

5. SUPPLY ELASTICITIES AND DEMAND SHOCKS ACROSS BOOMS

Our results point to a nationwide decline in housing supply elasticities. This suggests that aggregate demand shocks should have had a greater impact on house prices during the boom that started in mid-2012, whereas quantity should respond less.[14] To investigate this, we look at the impact of exogenous monetary policy shocks across the two booms.

13. 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.

14. In the Supplementary Appendix, we illustrate this point in a simple supply-demand model.

5.1 *High-Frequency Identification of Monetary Policy Shocks*

Our measure of monetary policy shocks is computed following the literature using high-frequency data to identify unexpected changes in the Fed policy rate (see, e.g., Gürkaynak, Sack, and Swanson 2005, Gertler and Karadi 2015, Nakamura and Steinsson 2018). 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 142 meetings over the two housing booms: 81 between 1997q1 and 2006q4 and 61 between 2012q3 and 2019q4.

We follow standard practice in the literature in transforming high-frequency data into low-frequency data (see, e.g., Gertler and Karadi 2015, Ottonello and Winberry 2020, Wong 2021). In particular, we first create a daily shock series by cumulating the daily surprises over the past 90 days. We then take quarterly averages of the cumulative daily shocks. Our quarterly shocks are characterized by roughly a 55%–45% distribution between expansionary and contractionary shocks over the full sample.[15] We note, however, that the HFI monetary policy shocks for the 2012q3–2019q4 period tend to be less volatile and smaller than for the 1997q1–2006q4 period, particularly during the zero lower bound period when the FOMC was actively targeting the longer end of the yield curve. This may pose identification challenges, which could result in less precise estimates for the effect of monetary policy shocks for the 2012q3–2019q4 period.

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. We use the 2-year Treasury bill yield as the relevant monetary policy indicator. The main reason for choosing a risk-free asset with a longer maturity than either the federal funds rate or the 1-year rate used by Gertler and Karadi (2015) is that the 2-year rate arguably better captures shocks to forward guidance about the future path of interest rates.

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, Schularick, and Taylor (2015), Ramey (2016), and Stock and Watson (2018) and use an IV local projection (LP) approach. The Jordà (2005) method offers some advantages over VAR models, since impulse re-

15. See Figure D.2 in the Supplementary Appendix for a visualization. The time-aggregation bias should not affect the results, as our quarterly shocks exhibit similar moments to the raw high-frequency data, see Table D.1 in the Supplementary Appendix.

sponses are less vulnerable to misspecification (Stock and Watson 2018). In addition, it easily accommodates nonlinearities, 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 2012q3–2019q4, by running a series of regressions for each horizon $h = 1,2\ldots,20$ 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, HPI, or in building permits, H, from period t to $t{+}h$.[16] MP_t is the monetary policy indicator (the 2-year Treasury bill yield), which is interacted with our estimated supply elasticities $El\hat{a}st_{i,t}$ 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).[17]

This large set of control variables helps minimize the potential 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. To account for the generated regressor issue caused by $El\hat{a}st_{i,t}$, we report bootstrapped standard errors based on 1,000 replications. We have also run a robustness check by clustering the standard errors at the state level. We find that the standard errors remain relatively similar to our baseline specification.

Our parameters of interest are $\beta^{Y,h}$ and $\gamma^{Y,h}$. From a theoretical point of view, one would expect an expansionary monetary policy shock to boost house prices ($-\beta^{HPI,h} > 0$), but that this effect becomes smaller for areas with higher housing sup-

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

17. Gilchrist and Zakrajšek (2012) define EBP 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.

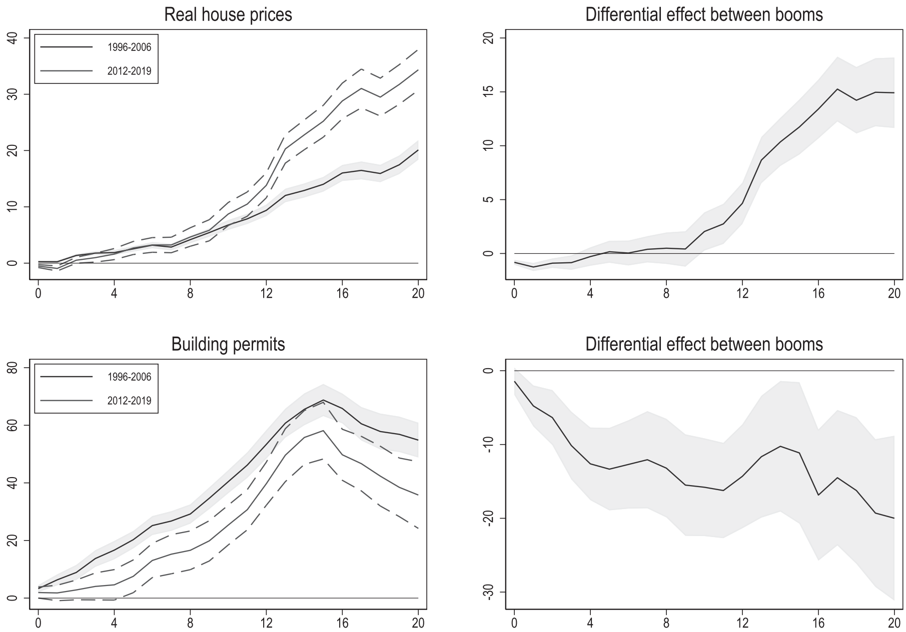

Fig 5. Responses to an Expansionary Monetary Policy Shock across Booms.

NOTES: Cumulative impulse responses to a 100 basis point decline in the 2-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 2012–19 and the 1996–2006 booms. The gray area and the dashed lines refer to 90% confidence bands produced by bootstrapping 1,000 replications.

ply elasticities ($-\gamma^{HPI,h} < 0$). 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 equation (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. We find that an expansionary monetary policy shock that lowers the 2-year Treasury 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 5).[18] Furthermore, we find that house prices rise by considerably more in the 2012–19 boom compared with the 1996–2006 boom. While price dynamics are similar in the short term, house prices in the most recent boom start to increase at a statistically significant faster pace after 2 years. For the same 100 basis points decline in government bond yields, the real house price response after 5 years is about twice as large for the 2012–19 boom com-

18. In Figure D.3 in the Supplementary Appendix, we show that results are robust to also using the 1-year Treasury bill yield as the monetary policy indicator.

pared to the previous boom. We estimate the opposite dynamics for building permits, which reacted more strongly to a monetary policy shock in the 1996–2006 boom. The finding that house prices respond more and quantity less to a demand shock during the most recent boom is consistent with our estimates in Section 2.5 that supply elasticities have declined.

5.3 *LP-IV: Sensitivity and Robustness*

Alternative monetary policy surprises. Since our monetary policy shocks tend to be less volatile and smaller for the 2012q3–2019q4 than the 1997q1–2006q4 period, we also study the robustness of our results to using alternative measures of monetary policy surprises that better capture the longer end of the yield curve. Swanson (2021) separately identify surprise changes in the federal funds rate, forward guidance, and large-scale asset purchases (LSAPs) for each FOMC announcement. Figure D.4 in the Supplementary Appendix shows that our results are robust to using these surprise series.

Sensitivity. To construct the quarterly surprise series, we first accumulate the daily series (by accumulating the daily surprises over the past 90 days) and then take quarterly averages. An alternative approach is to compute the quarterly surprises by taking simple averages of the daily surprises around FOMC meetings, as in Gazzani and Vicondoa (2020). Our results are robust to this alternative aggregation method and detailed results are displayed in Figure D.5 in the Supplementary Appendix.

6. CONCLUSION

We have provided evidence of a substantial and synchronized decline in local housing supply elasticities from the 1996–2006 housing boom to the housing boom that started in mid-2012. An implication of our finding is that the house price responsiveness to a given demand shock was higher during the most recent housing boom, 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 most recent housing boom than during the 1996–2006 boom. In contrast, we have found that the expansion in building permits was slightly smaller during the 2012–19 boom. Furthermore, our results point to significant heterogeneity in the responses across local housing markets. In particular, we estimate a substantially greater 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 regulatory constraints have 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 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 policymakers aimed at stimulating the real economy may have unintended effects by exacerbating the rise in house prices. Although it remains to be seen whether the decline in the elasticities is permanent or transitory, increasingly tighter land-use regulation suggests that it may be a permanent change. If this is a permanent change, monetary policymakers would need to take this into account.

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).

APPENDIX A: 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 onward 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, 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

TABLE A.1

Descriptive Statistics: Full Sample 1997q1–2019q4

	Obs	Mean	Std. Dev.	Min	Max
Real HPI (log)	23,178	4.8	0.2	4.0	5.5
House prices to income per capita	23,178	100.4	18.3	57.4	229.2
Building permits (log)	23,368	7.3	1.5	2.1	12.1
UNAVAL	23,368	0.3	0.2	0.0	0.9
WRLURI	23,368	−0.1	0.8	−1.8	4.3
Real personal income (log)	23,368	16.5	1.2	14.2	20.9
Real construction wages (log)	23,368	15.1	1.0	12.2	18.0
CPI (log)	23,368	5.3	0.2	4.1	5.7
Real mortgage originations (log)	23,367	13.8	1.4	10.0	19.1
Unemployment rate (%)	23,368	5.8	2.6	1.4	32.1
Population (log)	23,368	6.1	1.1	4.0	9.9
Population density (log)	20,320	5.4	0.9	1.8	7.9
Dependency ratio (%)	23,368	51.1	6.3	31.5	88.9
Black ratio (%)	23,157	11.7	11.2	0.2	54.1
Hispanic ratio (%)	23,157	11.6	14.9	0.4	92.5
Net domestic migration (%)	23,368	0.0	0.3	−12.5	3.1
Net international migration (%)	23,368	0.1	0.4	−43.1	15.1
Property crime rate (%)	18,956	8.1	0.4	1.1	9.0
State-level land-use regulation	504	2.5	1.4	0.4	7.3
ΔReal HPI (%)	23,178	0.3	1.9	−15.9	12.0
ΔReal personal income (%)	23,368	0.6	1.2	−9.3	12.2
ΔReal construction wages (%)	23,368	0.6	3.0	−20.3	19.1
ΔCPI (%)	23,368	0.5	0.7	−5.4	4.8
ΔReal mortgage originations (%)	23,366	0.6	18.7	−110.6	167.6
ΔUnemployment rate	23,368	0.0	0.4	−8.4	6.2
ΔPopulation (%)	23,368	0.2	0.5	−44.3	10.2
ΔPopulation density (%)	19,304	0.8	1.1	−28.7	6.1
ΔState-level land-use regulation	252	−5.6	29.2	−54.5	78.1

Note: The data sources are the Bureau of Economic Analysis, Bureau of Labor Statistics, Census Bureau, Federal Bureau of Investigation, Federal Housing Finance Agency, Gyourko, Saiz, and Summers (2008), Herkenhoff, Ohanian, and Prescott (2018), Home Mortgage Disclosure Act, Moody's Analytics, and Saiz (2010).

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, that is, 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, Saiz, and Summers (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.

TABLE A.2

Descriptive Statistics: 1997q1–2006q4 Housing Boom

	Obs	Mean	Std. Dev.	Min	Max
Real HPI (log)	10,160	4.8	0.2	4.1	5.5
House prices to income per capita	10,160	106.1	17.5	74.3	229.2
Building permits (log)	10,160	7.7	1.4	3.4	11.4
UNAVAL	10,160	0.3	0.2	0.0	0.9
WRLURI	10,160	−0.1	0.8	−1.8	4.3
Real personal income (log)	10,160	16.3	1.2	14.2	20.6
Real construction wages (log)	10,160	15.1	1.0	12.2	17.8
CPI (log)	10,160	5.2	0.1	4.1	5.5
Real mortgage originations (log)	10,159	14.0	1.5	10.0	19.1
Unemployment rate (%)	10,160	5.0	2.1	1.4	32.1
Population (log)	10,160	6.0	1.1	4.0	9.9
Population density (log)	7,112	5.3	0.9	1.8	7.9
Dependency ratio (%)	10,160	50.5	6.2	31.5	83.3
Black ratio (%)	10,108	11.4	11.2	0.2	50.5
Hispanic ratio (%)	10,108	9.7	14.0	0.4	89.7
Net domestic migration (%)	10,160	0.0	0.3	−12.5	3.1
Net international migration (%)	10,160	0.1	0.6	−43.1	15.1
Property crime rate (%)	7,584	8.3	0.4	1.1	9.0
State-level land-use regulation	252	2.6	1.4	0.7	7.3
ΔReal HPI (%)	10,160	0.9	1.4	−7.1	12.0
ΔReal personal income (%)	10,160	0.8	1.1	−9.3	12.2
ΔReal construction wages (%)	10,160	1.1	2.6	−12.3	17.3
ΔCPI (%)	10,160	0.6	0.6	−1.9	4.8
ΔReal mortgage originations (%)	10,158	1.8	21.0	−110.6	167.6
ΔUnemployment rate	10,160	0.0	0.4	−8.4	6.2
ΔPopulation (%)	10,160	0.3	0.6	−44.3	10.2
ΔPopulation density (%)	6,096	1.0	1.4	−28.7	6.1

Note: The data sources are the Bureau of Economic Analysis, Bureau of Labor Statistics, Census Bureau, Federal Bureau of Investigation, Federal Housing Finance Agency, Gyourko, Saiz, and Summers (2008), Herkenhoff, Ohanian, and Prescott (2018), Home Mortgage Disclosure Act, Moody's Analytics, and Saiz (2010).

Net domestic migration: the movement of people within the United States, computed as the difference between domestic inmigration and outmigration as a fraction of total population. Source: Census Bureau, and Moody's Analytics.

Net international migration: the movement of people in and out of the United States, computed as the difference between international inmigration and outmigration as a fraction of total population. Source: Census Bureau, and Moody's Analytics.

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.

TABLE A.3

Descriptive Statistics: 2012q3–2019q4 Housing Boom

	Obs	Mean	Std. Dev.	Min	Max
Real HPI (log)	7,430	4.8	0.2	4.0	5.5
House prices to income per capita	7,430	91.7	14.9	57.4	148.2
Building permits (log)	7,620	7.1	1.5	2.2	12.1
UNAVAL	7,620	0.3	0.2	0.0	0.9
WRLURI	7,620	−0.1	0.8	−1.8	4.3
Real personal income (log)	7,620	16.7	1.2	14.6	20.9
Real construction wages (log)	7,620	15.2	1.0	12.4	18.0
CPI (log)	7,620	5.5	0.1	4.5	5.7
Real mortgage originations (log)	7,620	13.6	1.4	10.5	18.2
Unemployment rate (%)	7,620	5.4	2.2	1.8	26.3
Population (log)	7,620	6.1	1.1	4.1	9.9
Population density (log)	7,620	5.4	0.9	2.0	7.9
Dependency ratio (%)	7,620	52.9	6.4	32.9	88.9
Black ratio (%)	7,476	12.0	11.2	0.7	54.1
Hispanic ratio (%)	7,476	13.6	15.7	0.8	92.5
Net domestic migration (%)	7,620	0.0	0.2	−1.9	1.0
Net international migration (%)	7,620	0.0	0.1	−0.4	0.9
Property crime rate (%)	6,442	7.8	0.3	6.6	8.7
State-level land-use regulation	252	2.5	1.5	0.4	6.2
ΔReal HPI (%)	7,430	0.8	1.5	−8.7	9.0
ΔReal personal income (%)	7,620	0.5	1.2	−8.1	6.6
ΔReal construction wages (%)	7,620	1.3	2.3	−20.3	19.1
ΔCPI (%)	7,620	0.3	0.4	−1.2	2.0
ΔReal mortgage originations (%)	7,620	1.1	13.0	−51.8	61.7
ΔUnemployment rate	7,620	−0.1	0.2	−2.5	1.6
ΔPopulation (%)	7,620	0.1	0.2	−1.4	1.0
ΔPopulation density (%)	7,620	0.6	0.9	−5.0	3.7

Note: The data sources are the Bureau of Economic Analysis, Bureau of Labor Statistics, Census Bureau, Federal Bureau of Investigation, Federal Housing Finance Agency, Gyourko, Saiz, and Summers (2008), Herkenhoff, Ohanian, and Prescott (2018), Home Mortgage Disclosure Act, Moody's Analytics, and Saiz (2010).

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 labor 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.

Land use regulation: model inferred land use regulation index. Data available at the state level. Source: Herkenhoff, Ohanian, and Prescott (2018).

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SUPPORTING INFORMATION

Additional supporting information may be found online in the Supporting Information section at the end of the article.

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