Journal of Public Economics 110 (2014) 82–100

Contents lists available at ScienceDirect

Journal of Public Economics

journal homepage: www.elsevier.com/locate/jpube

House prices and birth rates: The impact of the real estate market on the decision to have a baby ☆

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ARTICLE INFO

Article history:
Received 14 March 2012
Received in revised form 13 September 2013
Accepted 24 September 2013
Available online 17 October 2013

JEL Classification:
D10
J13
R21

Keywords:
Fertility
HPI
Credit constraints
Home equity
Cyclicality of births
MSA
Housing market

ABSTRACT

This project investigates how changes in Metropolitan Statistical Area (MSA)-level house prices affect household fertility decisions. Recognizing that housing is a major cost associated with child rearing, and assuming that children are normal goods, we hypothesize that an increase in house prices will have a negative price effect on current period fertility. This applies to both potential first-time homeowners and current homeowners who might upgrade to a bigger house with the addition of a child. On the other hand, for current homeowners, an increase in MSA-level house prices will increase home equity, leading to a positive effect on birth rates. Our results suggest that indeed, short-term increases in house prices lead to a decline in births among non-owners and a net increase among owners. The estimates imply that a $10,000 increase leads to a 5% increase in fertility rates among owners and a 2.4% decrease among non-owners. At the mean U.S. home ownership rate, these estimates imply that the net effect of a $10,000 increase in house prices is a 0.8% increase in current period fertility rates. Given underlying differences in home ownership rates, the predicted net effect of house price changes varies across demographic groups. In addition, we find that changes in house prices exert a larger effect on current period birth rates than do changes in unemployment rates.

1. Introduction

This project investigates how changes in Metropolitan Statistical Area (MSA)-level house prices affect household fertility decisions. The conceptual approach is based on an economic model of fertility that recognizes that changes in house prices potentially have offsetting effects on fertility. Assuming that children are normal goods, and recognizing that housing is a major cost associated with (additional) children, an increase in the price of housing will have a negative substitution effect on the demand for children in the current period, all else equal. This is true for both potential first-time homeowners (i.e., current non-owners who would buy a house with the addition of a child) and current homeowners who might buy a larger house with the addition of a child. On the other hand, for a homeowner, an increase in MSA-level house prices increases home equity. This could lead to an increase in birth rates among homeowners through two channels — a traditional wealth effect and/or an equity extraction effect. In either case, when house prices increase, homeowners might use some of their new housing equity to fund their childbearing goals. The net effect of house prices on aggregate birth rates will depend on individuals' responsiveness along these margins and rates of home ownership.

We are interested in identifying the causal relationship between movements in local area house prices and current period fertility rates. Conceptually, we are examining how short-term fluctuations in house prices affect current period fertility rates, separately for owners and non-owners, all else equal. Our main analyses focus on the housing price cycle of 1997 to 2006, a period of general housing price growth. We additionally separately consider the adjacent housing market cycles characterized by falling house prices. We begin our empirical investigation with a set of ordinary least square (OLS) regressions of MSA-demographic group-level fertility rates on MSA-level house prices interacted with a baseline measure of MSA-group-level home ownership rates, controlling for time-varying MSA conditions, and MSA fixed effects. To address the possibility that other local factors are biasing our OLS estimates we implement an instrumental variable (IV) strategy that exploits exogenous

☆ We thank Tom Davidoff, Seth Freedman, Phil Levine, Tim Moore, Kevin Mumford and seminar participants at the University of Maryland, Ohio State University, Washington University in St. Louis, Columbia University, and Harvard University for helpful comments on this project. We are also grateful to the editor of this journal and multiple anonymous referees for helpful suggestions. The analysis and conclusions set forth are those of the authors and do not indicate concurrence with other members of the research staff or the Board of Governors.

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0047-2727/$ – see front matter
http://dx.doi.org/10.1016/j.jpubeco.2013.09.009

variation in house price movements induced by variation across MSAs in their housing supply elasticity, as measured by Saiz (2010).

Both OLS and IV results indicate that as the proportion of individuals in a demographic cell who are home owners increases, an increase in house prices is conditionally associated with an increase in current period fertility rates. This is consistent with a positive "home equity effect" that dominates any negative price effect. The data also indicate that as the proportion of homeowners approaches zero, an increase in MSA-level house prices leads to a decrease in current period fertility rates, which is consistent with a negative price effect among non-owners. In general, the main results hold across race/ethnic groups and are equally driven by first, second, and higher-parity births.

These main results are statistically significant and economically meaningful. Employing our regression estimates in a straightforward simulation exercise, we find that a $10,000 increase in house prices is associated with a 5% increase in fertility rates in MSA cells with 100% ownership rates. For MSA cells with zero percent home ownership rates, we estimate a corresponding decrease in fertility rates of 2.4%. For an MSA-group, as the home ownership rates increase from 30 to 40%, the net effect of a $10,000 increase in house prices becomes positive. Under the assumption of linear effects, these estimates suggest that all else held constant, the roughly $108,000 average increase in house prices during the housing boom of 1997 to 2006 would have led to a 9% increase in births over that time.[1]

The main contribution of the paper is to provide an empirical examination of how aggregate movements in house prices affect aggregate level birth rates. First, as an issue of economic demography, it is informative to understand how movements in the real estate market affect current period birth rates, overall and for various demographic subgroups. Second, within the research literature on the nature of the demand for children, an examination of the effect of house prices on the fertility outcomes of homeowners constitutes a useful test of wealth effects. Third, our paper highlights the importance of including housing markets in any model of how economic conditions affect fertility outcomes. In fact, as an empirical matter, we find that changes in house prices exert a larger effect on current period birth rates than do changes in unemployment rates. Fourth, our results potentially speak to the role of credit constraints, and imperfect capital markets, in affecting the timing of fertility decisions. This is an issue that features prominently in the literature on the cyclicality of fertility timing, as reviewed in Hotz et al. (1997). Our finding of a positive effect among home owners suggests that some individuals may consume out of home equity to fund their childbearing goals. And finally, there is a literature on the tendency of individuals to consume out of housing wealth. To our knowledge, that literature has not previously considered children as a potential "consumption" good in this regard. Our results provide clear empirical support for the idea that house prices impact birth rates in a statistically significant and economically meaningful way.

2. Conceptual framework and related literature

There is a large literature in neoclassical economics investigating the nature and determinants of fertility in developed countries. In the simplest static approach to this question, parents are viewed as consumers who choose the quantity of children that maximizes their lifetime utility subject to the price of children and the budget constraint that they face. Children are conventionally thought to be normal goods, but an empirical puzzle presents itself in both time series and cross-sectional data, which tend to show a negative correlation between income and number of children.

There are two leading explanations for this observed correlation that maintain the basic premise of children as normal goods: (1) the quantity/quality trade-off (Becker, 1960) and (2) the cost of time hypothesis (Mincer, 1963; Becker, 1965). The first refers to the observation that parents have preferences for both the quantity and quality of children. If the income elasticity of demand for quality exceeds the income elasticity of demand for number of children, then as income rises, parents will substitute away from the number of children, toward quality per child. The second hypothesis attributes the observed negative relationship between income and fertility to the higher cost of parental time experienced by higher income families, either because of increased market wage rates or because higher household income raises the value of parental time in non-market activities. There is a long and active literature that attempts to estimate the effect of changes in family income and of own-prices on fertility.[2]

There exists a closely related literature investigating the cyclicality of fertility, which is a literature about fertility timing (e.g., Galbraith and Thomas, 1941; Becker, 1960; Silver, 1965; Ben-Porath, 1973). Changes in the unemployment rate are typically thought to affect the wages of women and their husbands. Under the standard assumption that women bear the primary responsibility for child rearing, it becomes optimal for woman to select into childbearing at times when their opportunity cost is lowest, that is, when economic conditions are least favorable. Another consideration affecting optimal timing with regard to unemployment rates is skill depreciation (Happel et al., 1984).[3]

In a world with imperfect capital markets and credit constraints, women might not be able to optimally time fertility with regard to opportunity cost and skill depreciation considerations. In particular, though some women might optimally choose to select into childbearing during economic downturns, they might not be able to afford to do this. Schaller (2011) provides a recent examination of this issue and explicitly considers the role of gender-specific labor market conditions. Her results confirm previous empirical findings that increase in overall unemployment rates are associated with decreases in birth rates. In support of the predictions of Becker's time cost model, she further finds that improved labor market conditions for men are associated with increases in fertility, while improved labor market conditions for women have the opposite-signed effect.[4]

Conceptually, the question of how real estate markets affect childbearing is more straightforward to consider because changes in house prices do not affect the cost of parental time. Our conceptual framework is thus not encumbered by considerations of skill depreciation or opportunity cost of time. We motivate our empirical model and interpret our estimated effects simply in terms of housing costs (which affect the price of childbearing) and housing income effects (which affect ability to consume in the current period). Our focus on current period prices and contemporaneous fertility allows us to look separately for price and "income" effects. Changes in the real estate market are expected to generate price effects because housing costs are estimated as the greatest portion of the annual cost of raising a child: greater than food, child care, or education (Lino, 2007).

[1] The population weighted average home price change for the 154 MSAs in our sample from 1997 to 2006 was $108,038.

[2] The key empirical challenge in this literature is to find variation in family income or the price of children that is exogenous to women's (or couple's) preferences and the opportunity cost of women's time. Many of these papers are reduced-form in nature, and include examinations, for example, of the effect of direct pro-natalist government payments (e.g., Milligan, 2005; Cohen et al., 2007) and of exogenous changes in income (Lindo, 2010; Black et al., 2011).

[3] There exists a class of dynamic or life-cycle models of fertility decisions, which recognize that changes in prices and income over the life cycle may result in changes in the timing of childbearing, even if they do not cause completed lifetime fertility to change. The Handbook chapter by Hotz et al. (1997) provides an overview of these theoretical models. Heckman and Walker (1990) provide an empirical examination of the effect of income and wages on life-cycle fertility using data from Sweden.

[4] Dehejia and Lleras-Muney (2004) suggest that relatively more white women opt into childbearing during economic downturns than black women; they attribute this difference to credit constraints facing blacks. Neither Schaller (2011) nor we find evidence in the data consistent with this idea. In particular, we find a statistically significant negative relationship between unemployment rates and birth rates among whites and a statistically insignificant relationship among blacks.

We qualify the term “income” when we talk about housing income effects because an increase in house prices does not necessarily imply increased wealth or income for home owners. If price increases are viewed to be permanent and homeowners view their home as a store of wealth, an increase in house prices can be thought of as an increase in (perceived) wealth for existing homeowners. This could lead to an increase in the demand for children in the current period, as well as in a completed lifetime setting. But, if homeowners do not intend to “cash out” and move to a lower-priced real estate market during their lifetime, or if they view the increase in house prices as transitory and expect it to be undone at a later period, there is no change in actual wealth or permanent income. However, if homeowners are otherwise credit constrained but can liquefy increases in home equity, there can be an increase in current period accessible income and this could lead to an increase in current period birth rates. This may or may not lead to an increase in completed lifetime fertility. For the sake of convenience of exposition, we refer to this general class of explanations as a “home equity effect”.

There is a large body of research on the propensity for households to fund current consumption out of housing wealth.[5] This literature recognizes the two distinct effects of housing values on consumption: the traditional wealth effect and a home equity extraction effect. A recent paper by Mian and Sufi (2009) estimates that the average homeowner extracted 25 to 30 cents for every dollar increase in home equity during the 1997 to 2009 period. They further find that money extracted from increased home equity was not used to purchase new real estate or pay down high credit card balances, which they interpret as suggesting that borrowed funds were used for consumption or home improvement expenses. In addition, they find that home equity-based borrowing was strongest among younger households. These findings allow for the possibility that during the recent housing boom, individuals and couples used some of their increased housing equity to fund child-related expenses.

One could reasonably argue that in contrast to unemployment rates – which are generally understood to be cyclical – movements in the housing market over the period we analyze were likely to have been perceived at least in part as permanent. This would follow from the observation that the national trend in housing prices between 1997 and 2006 was steadily increasing. This suggests our results may be indicative of a change in completed fertility, as opposed to simply a story about timing or cyclicality. We give a cursory treatment of this possibility in our empirical analyses below – in particular by looking at higher-order births – but we leave it to future research to thoroughly examine this possibility.

Finally, we acknowledge that we talk about fertility throughout the paper as though it is a simple decision. Of course, fertility is a stochastic outcome, albeit one that is to a large extent controllable by individual's actions with regard to sexual activity, contraceptive use, fertility treatments, and abortion. We recognize, however, that latent demand for fertility timing will not be perfectly realized. Thus, any response we see of fertility to house prices will be a muted reflection of a couple's desired fertility response.

3. Data and empirical approach

The main empirical approach of this paper is to empirically relate MSA-level fertility rates to demeaned MSA-level house prices, interacting house prices with a baseline measure of group-level home ownership rates and controlling for time-varying MSA-level characteristics. The three main data requirements are (1) MSA-level fertility rates, (2) MSA-level house prices, and (3) group-level home ownership rates. In this section we describe our main data sources and briefly describe how we construct the relevant variables. Table 1 provides details on explanatory variables and associated data sources.

[5] See for example, Case et al. (2005), Benjamin et al. (2004), Bostic et al. (2009), Haurin and Rosenthal (2005).

3.1. Data

Data on births come from the Vital Statistics Natality Files, years 1990 to 2007. Vital statistics data contain birth certificate information for virtually every live birth that takes place in the United States. Vital statistics data identifies the race/ethnicity, marital status, age, and education of the mother, as well as some limited information about the pregnancy conditions and the baby's health status at time of birth. For the purposes of matching births to our explanatory variables, we create a file of conceptions for the years 1990 to 2006, using information on the date of birth and length of gestation to identify year of conception. We do this because in terms of the decision-making process, the most relevant decision is the decision to get pregnant in a given time period. It is thus the economic conditions that exist at the conception decision point that are relevant, as opposed to the economic conditions in place at the time when the birth actually occurs (typically 40 weeks later.) To be precise, our analysis sample is a sample of conceptions that result in live births in year *t*.

We construct MSA-year-group-level fertility rates by aggregating births and female population counts to the MSA-year-group cell, where groups are defined by the interaction of race/ethnicity and age category. We define three mutually-exclusive race/ethnic groups: Non-Hispanic White, Non-Hispanic Black, and Hispanic. We exclude other race/ethnicities from the analysis. We define two age categories, 20–29 and 30–44. We obtained annual female population counts (by age, race, ethnicity, and county) from the National Center for Health Statistics (2003, 2010). We use these data to construct MSA-group-level fertility rates, defined as the total number of births to women in the MSA-year-group cell divided by the MSA-year-group population. We obtained access to confidential natality files that identify the mother's state and county of residence. We use the county-level identifiers in the confidential Vital Statistics Natality Files to construct MSA-level fertility rates, using the MSA definitions that are used in the federal housing datasets: 5-digit MSAs and Divisions as defined by the Office of Management and Budget in December 2009 (Bulletin 10-02).

We identify a total of 384 MSAs in the birth records. We restrict our sample to MSAs that have at least five births in every year-group cell, which leaves us with a sample of 222 MSAs. When we further restrict the sample to those MSAs for which all explanatory variables used in the baseline specification are available, we are left with a sample of 154 MSAs.[6]

The main data source used to construct MSA-level house prices is the Federal Housing Finance Agency (FHFA) housing price index (HPI), previously known as the OFHEO housing price index. The FHFA index is available for nearly all metropolitan areas in the United States.[7] It measures the movement of single family home prices by looking at repeat mortgage transactions on homes with conforming, conventional mortgages purchased or securitized through Fannie Mae or Freddie Mac since 1975.[8] Since the index looks at repeat mortgages of the same home, it is continually revised to reflect current MSA

[6] This process eliminates 60% of MSAs, but only about 15% of births.

[7] FHFA requires a metro area to have at least 1000 transactions before it is published.

[8] Conventional mortgages are those that are neither insured nor guaranteed by the FHA, VA, or other federal government entities. Mortgages on properties financed by government-insured loans, such as FHA or VA mortgages, are excluded from the HPI, as are properties with mortgages whose principal amount exceeds the conforming loan limit. Mortgage transactions on condominiums, cooperatives, multi-unit properties, and planned unit developments are also excluded. This contrasts to the widely used alternative Case–Shiller index, which includes all homes, but is only available for 37 states and a more limited set of MSAs. Additional differences between the two indices are that the Case–Shiller index puts more weight on more expensive homes and the Case–Shiller index uses purchases only, whereas the FHFA index also includes refinance appraisals. As a robustness check, we have re-estimated our results using the Case–Shiller index. Note that we use the “all transactions” version of the FHFA index, which includes both sales and refinancings of existing mortgages. We do not use the “sales only” version of the index because it is available for only a small subset of MSAs.

Table 1
Aggregate variables

Variable	Mean	Standard deviation	Source	Description	Geographic detail
House Price Index (HPI)	163.64	38.14	Federal Housing Finance Agency (FHFA)	House Price Index (all transactions)	MSA Divisions (2009)
Median home price	$162,356	$89,700	2000 Census and FHFA	2000 Median home price scaled by annual mean FHFA HPI	County, MSA Divisions (2009)
Male wages:					
25th percentile	$12.94	$2.33	Current Population Survey (CPS)	Individual wage and salary income divided by the product of weeks and hours worked for adult men working full time	Primary MSAs (1983, 1993, 2003)
50th percentile	$19.10	$3.62			
75th percentile	$28.09	$6.52			
All wages:					
25th percentile	$11.58	$1.79	Current Population Survey (CPS)	Individual wage and salary income divided by the product of weeks and hours worked for adult men working full time	Primary MSAs (1983, 1993, 2003)
50th percentile	$17.00	$2.71			
75th percentile	$24.91	$4.27			
Mean wage	$23.17	$4.38	Current Population Survey (CPS)	Mean of Individual wage and salary income divided by the product of weeks and hours worked for all adults working full time	Primary MSAs (1983, 1993, 2003)
Unemployment rate	4.34	1.74	Bureau of Labor Statistics	Mean annual unemployment rates from Local Area Unemployment Statistics	County
Income per capita	$39,740	$7,149	Bureau of Economic Analysis	Sum of income from all sources divided by total population from Regional Economic Accounts	County
Average rent	$785	$214	Department of Housing and Urban Development	Mean fair market rent for 0-4 bedroom residences	County
Housing supply elasticity	1.96	0.97	Saiz (2010)	Measure of housing supply elasticity	Primary MSAs and NECMAs (2009)
Fraction college	0.19	0.19	Current Population Survey (CPS)	Fraction of MSA-group cell with a college education	Primary MSAs (1983, 1993, 2003)

Notes: Listed are aggregate level variables and their means for the 154 MSAs in our sample. All variables are aggregated up the MSA level from the level of geographic detail (column 6) available using the crosswalk procedure described in the text and data Appendix A. All nominal values are CPI adjusted to 2006 dollars.

boundaries. This is the reason we must use the most current definitions of MSAs in constructing the birth data. We annualize the index (which is available quarterly) by taking the mean value of the index over the four quarters of a year.

We use the FHFA index to construct real house prices for each MSA-year by combining it with information on median home values obtained from the 2000 census. The 2000 Census records median home values for each county in the U.S. We use the same county crosswalk used to construct MSAs in the birth data to construct MSA-level median 2000 house values, which are the population-weighted average across all counties in each MSA. Home values are scaled by the relevant change in the FHFA index over time and are adjusted to 2006 dollars using the CPI-U "All items less shelter" series. This measure serves as a proxy for real house price movements of median value homes in each MSA.[9]

The third main variable we need to construct is a measure of mean group-level home ownership rates at the MSA level. This is key to our analysis because conceptually, we expect that there would be heterogeneous responses of birth rates to home prices across groups with different rates of home ownership. The Vital Statistics data do not include information about home ownership status, so we cannot separately tabulate current period births (or conceptions) for home owners and non owners. Furthermore, we ideally do not want to use an individual-level measure of realized home ownership rate, because it is potentially endogenously determined with childbearing outcomes. Our implemented solution is to use MSA-group-level home ownership rates calculated from the 1990 five percent sample of the decennial census. As mentioned above, groups are defined by race/ethnicity and age category. We match the MSA definitions provided in the Census to the 2009 MSA definitions used for the birth and housing price data according to the crosswalk procedure described in the Appendix A. To be clear, our group-level measure of home ownership is taken at baseline and is time invariant.

3.2. Descriptive statistics and trends

Fig. 1 displays trends in mean (CPI adjusted) house prices, constructed as described above, in our sample, both in levels and yearly percentage changes between year *t-1* and *t*. Fig. 1 also displays house prices alternatively constructed using the Case–Shiller index to scale 2000 median home prices. The three housing cycles that fall within our period of study are highlighted: the 1990–1996 period of price decline, the 1997–2006 housing boom, and the subsequent 2007–2010 housing bust. Appendix Table 1 lists the 154 MSAs included in our analysis sample, ranked according to the percentage increase in housing prices between 1997 and 2006.[10] Fig. 2 displays the time-series correlation between fertility rates and house prices and then between fertility rates and unemployment rates, for the period 1990–2006, averaged across the MSAs in our sample. These plots suggest that movements in fertility rates track movements in house prices fairly closely, particularly in

[9] We adopt this procedure from Glaeser et al. (2008).

[10] There is an active literature exploring various explanations for the boom and bust in house prices experienced in recent decades. As summarized by Sinai (2012) – who offers citations for the various factors – these potential explanations include "changing interest rates, subprime lending, irrational exuberance on the part of home buyers, a shift to speculative investment in housing, contagion and fads, and international capital flows." Sinai's 2012 paper presents a set of empirical facts about the recent housing cycle, including information about how the amplitude and timing of house price appreciation and depreciation varied across MSAs. One of the observations he makes that is particularly relevant to our current empirical approach is that "demand fundamentals" do not have the same amplitude as price cycles nor does the time pattern of the growth in fundamentals match the timing of the growth in house prices across MSAs.

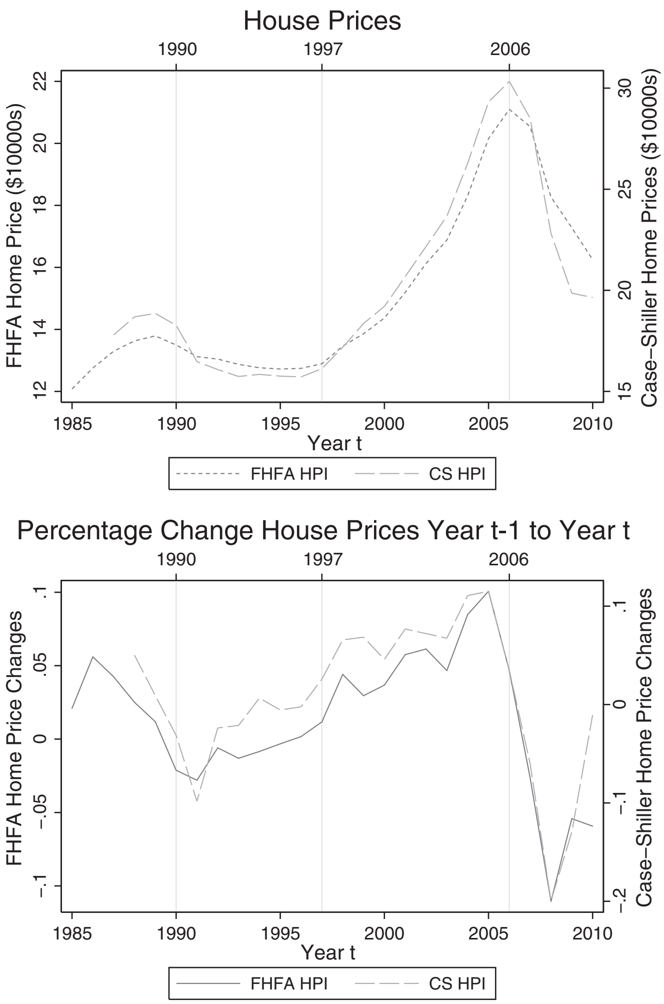

Fig. 1. Housing price index (FHFA and Case–Shiller). Notes: House prices are calculated using 2000 MSA median home values, which are scaled by either the FHFA house price Index or the Case–Shiller house price index to create MSA-year median home values, which are then averaged over the 154 MSAs (27 MSAs for the Case–Shiller Index) in our sample each year 1984–2010. Both are adjusted to 2006 dollars using CPI-U "all items less shelter" series. Percentage change in home prices is calculated as $(HousePrice_t - HousePrice_{t-1})/HousePrice_{t-1}$. In both figures, the left y-axis represents the mean value of the FHFA-constructed prices and the right y-axis represents the mean value of the Case–Shiller constructed prices.

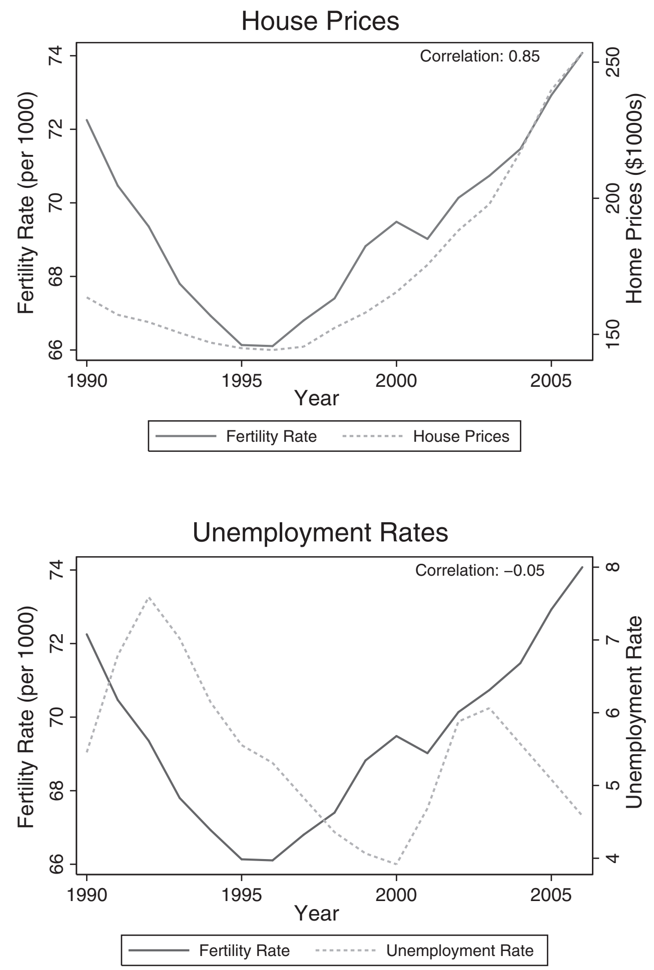

Fig. 2. Fertility rates and macro indicators. Notes: Displayed are trends in fertility rates, housing prices, and unemployment rates. Annual fertility rates (births per 1000 women) are calculated using yearly totals of MSA-level births to women aged 20–44 divided by total female population aged 20–44, both obtained from the National Center for Health Statistics National Vital Statistics System. House prices are 2000 median home values scaled by the Federal Housing Finance Agency (FHFA) housing price index, and are displayed in 2006 dollars. Unemployment rate is the annual mean unemployment rates taken from Bureau of Labor Statistics local area unemployment statistics. All three measures are yearly mean values calculated based on the 154 MSAs in our sample.

more recent periods. In fact, a comparison of the graphs reveals that the time-series correlation between aggregate fertility rates and housing prices is much greater than it is between aggregate fertility rates and unemployment rates, .85 versus −.04. This provides a *prima facie* case for the importance of considering housing prices when investigating how economic conditions affect current period birth rates.

Table 2 provides summary statistics from the 1997 to 2006 Vital Statistics Natality Files and the 1990 and 2000 Census. These data are used collectively in various analyses presented below. All measures are female-population weighted. The first three columns summarize the main dependent variable of interest: fertility rates (group-level births per 1000 women age 20–44), overall and for first and higher parity births. The overall fertility rate in our sample is 70 births per 1000 women aged 20–44. The highest fertility rates are found among Hispanics aged 20–29: 154 births per 1000 women. The lowest rate is among Blacks aged 30–44: 38 births per 1000 women.

The next column summarizes data from the 1990 census on MSA-group-level home ownership rates. The overall home ownership rate among our sample of women aged 20–44 is 44%. The highest home-ownership rates are found among older (age 30–44) White women, who have an ownership rate of 67%. The lowest rates are found among younger (age 20–29) Black women, whose ownership rate is on average 8%. This indicates that there is substantial variation across groups in rates of ownership. For the sake of comparison, the next column shows the rates as calculated from the 2000 census. Comparing the group-level ownership rates in 1990 and 2000 we see that home ownership rates are extremely stable over this time period. The final column displays the range of the 1990 ownership rate across MSAs, within each demographic group. These numbers indicate that in addition to the substantial variation across groups in rates of home ownership, there is also substantial variation within groups across MSAs.

3.3. Empirical specification

Our initial empirical analysis consists of ordinary-least squares regressions (OLS) at the MSA-group-year level. For our baseline analysis, we restrict our attention to the housing cycle of 1997–2006. This facilitates interpretation as the period was one of nearly uniform house price growth, and is recognized by the real estate literature as a housing boom period. We will subsequently consider two housing bust periods: the early 1990s bust period (1990–1996) and the post-2006 housing bust (2007–2010).

Table 2
Summary statistics.

	Vital statistics			Census		
	Fertility rate (1000)	First birth fertility rate	Higher birth fertility rate	Home ownership rate 1990	Home ownership rate 2000	Min/max ownership rate
All	70.09	24.60	45.49	0.44	0.44	[0.00, 0.80]
	(36.40)	(16.55)	(22.68)	(0.23)	(0.23)	
White	88.48	42.02	46.46	0.27	0.25	[0.10,0.39]
20–29	(20.76)	(6.96)	(14.43)	(0.06)	(0.07)	
Black	118.00	40.19	77.81	0.08	0.10	[0.00,0.29]
20–29	(17.18)	(5.11)	(16.78)	(0.03)	(0.03)	
Hispanic	153.78	54.70	99.09	0.14	0.15	[0.00,0.50]
20–29	(31.03)	(9.82)	(24.71)	(0.06)	(0.06)	
White	48.28	14.84	33.44	0.67	0.68	[0.47,0.80]
30–44	(8.30)	(4.62)	(4.74)	(0.07)	(0.08)	
Black	37.60	7.90	29.70	0.34	0.35	[0.06,0.60]
30–44	(7.93)	(2.75)	(5.80)	(0.08)	(0.08)	
Hispanic	59.69	10.81	48.88	0.40	0.41	[0.00,0.80]
30–44	(11.19)	(3.07)	(10.69)	(0.13)	(0.13)	

Notes: Fertility rates are total births over the total female population in each MSA, year of conception, age category and race/ethnicity cell for women aged 20–44. Mean home ownership rates are calculated in 1990 Census by year, MSA, age category, and race/ethnicity. Min/max home ownership rates are based on 1990 data. Sources for aggregate birth data and population data are Vital Statistics birth certificate data (1997–2007) and population data (1996–2006), and for home ownership data is the decennial Census (1990 and 2000). All means displayed are population weighted. Standard deviations are in parentheses.

We estimate regression models of the following form:

$$\begin{aligned} ln\left(FertRate_{mtg}\right) = {} & \beta_0 + \beta_1\left(HousePrices_{mt-1} * OwnRate_{mg}\right) \\ & + \beta_2 HousePrices_{mt-1} + \beta_3 OwnRate_{mg} + \beta_4 \mathbf{X}_{mt-1} \\ & + FracColl_{mgt-1} + \gamma_m + \gamma_t + \gamma_g + \gamma_m * (t-1) \\ & + \gamma_m * OwnCat_{mg} * (t-1) + \epsilon_{mgt}. \end{aligned} \tag{1}$$

The level of analysis is an MSA-year-group cell. In the above equation, the subscript m denotes MSA, t denotes year of the birth (where t-1 refers to the year of conception), and g denotes group.[11] There are six groups, defined by the interaction of our three race/ethnic groups (Non-Hispanic White, Non-Hispanic Black, and Hispanic) and two age categories (age 20–29 and age 30–44). Our final analysis sample consists of 9240 observations (10 years $*$ 6 groups $*$ 154 MSAs). All regressions are weighted by the total number of births in each cell.[12]

The coefficients of primary interest are β_1 and β_2, which capture the conditional effect, respectively, of MSA-year house price index (HPI) interacted with a baseline measure of MSA-group-level ownership rates and the conditional main effect of the MSA-year house prices ($HousePrice_{mt-1}$) on fertility rates. The former indicates how an increase in home ownership rates affects the relationship between de-meaned (and sometimes de-trended) MSA house prices and births. The conditional main effect of $HousePrice_{mt-1}$ indicates how movements in house prices affect fertility rates net of ownership interactions, all else held constant. We interpret this to be the conditional relationship between $HousePrice_{mt-1}$ and log fertility rates among a non-home-owning population of households.

The variable $OwnRate_{mg}$ is the MSA-group-level home ownership rate measured in the 1990 five percent sample of the decennial census. This measure is taken at baseline to minimize concerns about the endogeneity of year-specific MSA home ownership rates and year-specific MSA fertility rates. However, home ownership rates are quite stable over time within groups, which means that the baseline measure is highly predictive of current period home ownership rates. Therefore, this approach does not entirely eliminate any concern about endogenously determined current period births and our measure of home ownership rates. We control for this conditional main effect to facilitate a causal interpretation of β_1, but we are careful not to assign a causal interpretation to the coefficient on ownership rates.

We are interested in identifying the causal relationship between lagged house prices and fertility rates. It is thus important to control for other time-varying MSA-level economic conditions that potentially covary with real estate markets and also fertility timing decisions. Our regression specification includes controls for MSA-year unemployment rate, MSA-year male wages included in the vector X_{mt} in Eq. (1). The specification also controls for $FracColl_{mgt}$, the fraction college educated in each MSA-group-year. This is calculated as a three year moving average using data from the Current Population Survey. Data on MSA-year level unemployment rates come from the Bureau of Labor Statistics (BLS) Local Area Unemployment Statistics. Our measure of MSA-year level male wages is the 25th, 50th, and 75th percentile male wage, which was calculated by MSA and year in the Current Population Survey. Percentiles of the wage distribution were constructed based on hourly earnings for full-time, full-year male workers.[13] Unemployment rates were collected at the county level and aggregated to MSAs using the crosswalk procedure described in the Appendix A. The wage and fraction college measures were calculated using the MSA definitions available in the CPS and translated to 2009 MSAD definitions using the crosswalk procedure.

The regression also includes controls for MSA fixed effects (γ_m), year fixed effects (γ_t), group fixed effects (γ_g), and in some specifications, MSA-specific time trends ($\gamma_m * (t-1)$) and MSA-ownership-cell-specific time trends ($\gamma_m * OwnCat_{mg} * (t-1)$). It is imperative that the regression specification control for MSA fixed effects so that the estimated relationship between house prices and birth rates is not confounded by time-invariant differences in preferences for children across MSAs. If couples with lower preferences for children sort into areas with higher costs of living – driven by other amenities – there will be a negative correlation between house prices and fertility.[14] Given our goals in this paper, we want to isolate the effect of house prices on current period fertility net of these sorting patterns. The regression estimate of the relationship between house prices and birth rates is identified off within-MSA changes in house prices. We additionally include MSA-specific trends and MSA-ownership category-specific trends in the model to allow for the possibility that individuals with plans to expand their families choose to locate in MSAs with upward or downward trending prices, and that owners and renters may behave differently in this respect.

[11] For the sake of convenience, we write t-1, but our empirical analysis is precise in dating the year of conception by taking the date of birth and subtracting off the reported weeks of gestation.

[12] Results alternatively weighting by total female population in each cell are similar and available from the authors upon request.

[13] We construct wages as in Autor et al. (2008). We define full time as 35 or more hours per work, and full year as 40 or more weeks worked in the past year. We drop individuals who make less than one half the 2006 minimum wage (in 2006 dollars). Top-coded observations are multiplied by 1.5.

[14] For example, consider the hypothetical case of two couples, in which one moves to San Francisco, where household expenses are high, because they expect to have few children and spend their time and money instead indulging in city-type amenities. The other couple moves to Wichita, in expectation of buying a big house at a much lower cost per square foot, and filling it with kids. If these couples are typical, then high-latent-fertility couples will sort into lower priced real estate markets and low-latent-fertility couples will sort into lower priced real estate. Simon and Tamura (2008) examine the cross-sectional relationship between fertility and the price of living space across U.S. metropolitan areas, as captured by the average rent per room in an urban area (calculated among renting households.) Their baseline specification, which controls for region effects and demographic composition, suggests that a one percent increase in rent is associated with 0.16 fewer children per household.

4. Estimation results

4.1. Ordinary least squares specifications

Table 3 presents the results of estimating Eq. (1). Column 1 reports the results with all fixed effects included, but without MSA-specific controls for labor market conditions. This sparse specification yields a point estimate of β_1 of 0.0468 and a point estimate on β_2 of -0.0124, both statistically significant at the one percent level. The positive and statistically significant point estimate on the interaction term $HousePrice_{mt-1} * OwnRate_{mg}$ indicates that as home ownership rates increase, higher house prices lead to an increase in current period births, all else held constant. This implies that a positive home equity effect dominates any negative price effect among current home owners. The negative and statistically significant point estimate on $HousePrice_{mt-1}$ is consistent with a negative price effect of house prices on current period fertility for non-home owners. Column 2 adds the unemployment rate and wage measures. The main point estimates of interest are qualitatively unchanged. Looking at other explanatory variables, we see that the estimated coefficient on the mean ownership rate is positive, but statistically insignificant. As noted above, we do not propose a causal interpretation to this relationship. The unemployment rate is found to be negatively related to the fertility rate in all specifications, but it does not enter with statistical significance.

Next, we include MSA-specific time trends in the model to allow for the possibility that individuals with plans to increase or decrease their fertility move into MSAs with upward or downward trending house prices. Columns 3 and 4 report the results with MSA specific linear trends and MSA specific quadratic trends, respectively. The pattern remains the same – a positive coefficient on $HousePrice_{mt-1} * OwnRate_{mg}$ and a negative coefficient on $HousePrice_{mt-1}$ – and the magnitudes of the coefficients are similar to the specification without any trend terms included, giving us no reason to suspect that individuals with plans to increase or decrease their fertility systematically move into MSAs with upward or downward trending house prices.

If there exist trends of this kind that are distinct for groups with high and low ownership rates, the estimated β_1 might be a biased estimate of the conditional causal effect of interest. We thus additionally include in the model separate MSA specific time trends based on whether a group's level of ownership (in a particular MSA) is above or below the median home ownership rate (of 30.5% in our sample of MSA*group cells), yielding two values of $OwnCat_{mg}$. These trends allow, for example, White women aged 30–44 in the Boston metro area to be on a different trend then Black women aged 20–29 in the Boston metro area. Column 5 displays the results estimating the model with distinct MSA-specific trends for cells with ownership rates above and below the median home ownership rate. These trends also do not alter the pattern of estimates, but the point estimates are attenuated toward zero. In this model, the estimated coefficient on the $HousePrice_{mt-1} * OwnRate_{mg}$ interaction is 0.0276 (with a standard error of .00375) and the estimated coefficient on $HousePrice_{mt-1}$ is -0.00509 (standard error of .00184). This is arguably the most conservative of the OLS specifications. These estimates suggest that if house prices increase by $10,000, as we move

Table 3
Housing prices and fertility rates: 1997–2006.

Dep. Var. $\log(FertRate)_{mgt}$	(1)	(2)	(3)	(4)	(5)
$HousePrice_{mt-1} * OwnRate_{mg}$	0.0468***	0.0468***	0.0481***	0.0485***	0.0276***
	(0.00443)	(0.00448)	(0.00481)	(0.00488)	(0.00375)
$HousePrice_{mt-1}$	−0.0124***	−0.0128***	−0.0111***	−0.0160***	−0.00509***
	(0.00103)	(0.00115)	(0.00243)	(0.00195)	(0.00184)
$OwnRate_{mg}$	0.0545	0.0542	0.0460	0.0445	0.0852
	(0.318)	(0.318)	(0.320)	(0.323)	(0.267)
WhiteAge 20–29	0.948***	0.949***	0.953***	0.953***	0.902***
	(0.128)	(0.128)	(0.127)	(0.128)	(0.120)
BlackAge 20–29	1.335***	1.334***	1.337***	1.335***	1.255***
	(0.185)	(0.185)	(0.184)	(0.186)	(0.174)
HispanicAge 20–29	1.569***	1.569***	1.570***	1.567***	1.500***
	(0.164)	(0.164)	(0.163)	(0.164)	(0.164)
BlackAge 30–44	0.0112	0.0111	0.0121	0.00995	−0.0484
	(0.0749)	(0.0748)	(0.0743)	(0.0749)	(0.0680)
HispanicAge 30–44	0.371***	0.371***	0.369***	0.365***	0.328***
	(0.0790)	(0.0790)	(0.0784)	(0.0791)	(0.0599)
$FracColl_{mgt-1}$	−0.349***	−0.352***	−0.375***	−0.395***	−0.358***
	(0.0641)	(0.0648)	(0.0682)	(0.0710)	(0.0679)
$UnempRate_{mt-1}$		−0.00255	−0.00116	−0.00156	−0.00155
		(0.00335)	(0.00211)	(0.00204)	(0.00200)
$25thWage_{mt-1}$		0.00138	0.000674	0.000237	0.000744
		(0.00106)	(0.000660)	(0.000550)	(0.000621)
$50thWage_{mt-1}$		0.000166	0.000978*	0.000475	0.000839*
		(0.000904)	(0.000534)	(0.000424)	(0.000480)
$75thWage_{mt-1}$		0.000623	0.000241	0.000182	0.000199
		(0.000552)	(0.000236)	(0.000216)	(0.000230)
MSA fixed effects	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes
MSA trends	No	No	Yes	Yes	Yes
MSA quadratic	No	No	No	Yes	No
MSA-own category trends	No	No	No	No	Yes
R^2	0.908	0.908	0.910	0.910	0.937
Number of MSAs	154	154	154	154	154
N	9240	9240	9240	9240	9240

Notes: Fertility rates are total births over the total female population in each MSA, year of conception, age category and race/ethnicity cell for women aged 20–44. Mean home ownership rates are calculated in 1990 Census by year, MSA, age category, and race/ethnicity. Fraction of cell that is a college graduate is matched by MSA, year, age category and race. House prices (10,000 s), unemployment rates, and male wages are matched by MSA and year of conception. Data sources are: Vital Statistics (births, population), Census and Federal Housing Finance Agency (house prices), Current Population Survey (wages, fraction college), and Bureau of Labor Statistics (unemployment rates). All specifications are weighted by the total number of births in the cell. Standard errors adjusted for clustering at the MSA level are in parentheses. * $p < 1$, ** $p < .05$, *** $p < .01$.

Table 4
Housing prices and fertility rates by MSA supply elasticity: 1997–2006.

	(1)	(2)	(3)	(4)	(5)
	Low elasticity	High elasticity	First stage 1: Hprice ∗ Own	First stage 2: Hprice	IV
$HousePrice_{mt-1} * OwnRate_{mg}$	0.0428***	0.0609***			0.0723***
	(0.00448)	(0.0128)			(0.00997)
$HousePrice_{mt-1}$	−0.0104***	−0.0170***			−0.0239***
	(0.00117)	(0.00498)			(0.00328)
$Elasticity_m * HPI_t * OwnRate_{mg}$			−0.0137***	−0.000647**	
			(0.00200)	(0.000287)	
$Elasticity_m * HPI_t$			−0.0111***	−0.0571***	
			(0.00186)	(0.00956)	
$OwnRate_{mg}$	0.143	−0.394	17.93***	0.193	−0.171
	(0.433)	(0.363)	(2.332)	(0.353)	(0.291)
$FracColl_{mgt-1}$	−0.419***	−0.111	4.995***	0.674	−0.492***
	(0.0856)	(0.0762)	(0.804)	(0.673)	(0.102)
$UnempRate_{mt-1}$	−0.00644	0.0000946	−0.111	−0.994***	−0.0137**
	(0.00459)	(0.00270)	(0.0887)	(0.321)	(0.00563)
$25thWage_{mt-1}$	0.00253	−0.00106	−0.0147	0.0338	0.00226
	(0.00152)	(0.000871)	(0.0319)	(0.101)	(0.00166)
$50thWage_{mt-1}$	0.00134	−0.00117*	0.0154	0.116	0.00153
	(0.00141)	(0.000674)	(0.0268)	(0.0786)	(0.00115)
$75thWage_{mt-1}$	0.00111	0.000417	−0.0000601	0.0304	0.00120*
	(0.000790)	(0.000314)	(0.0188)	(0.0434)	(0.000704)
MSA fixed effects	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes
R^2	0.900	0.939	0.876	0.932	0.897
F Statistic			29.60	18.32	
Number of MSAs	77	77	154	154	154
N	4620	4620	9240	9240	9240

Notes: Fertility rates are total births over the total female population in each MSA, year of conception, age category and race/ethnicity cell for women aged 20–44. Mean home ownership rates are calculated in 1990 Census by year, MSA, age category, and race/ethnicity. Fraction of cell that is a college graduate is matched by MSA, year, age category and race. House prices (10,000 s) are matched by MSA and year of conception (or years prior to conception where noted). Elasticity refers to the Saiz (2010) supply elasticity measure and HPI refers to the national version of the FHFA house price index. First stage 1 refers to the first stage regression where the dependent variable is $HousePrice_{mt-1} * OwnRate_{mg}$ and first stage 2 refers to the first stage regression where the dependent variable is $HousePrice_{mt-1}$. All regressions include group, MSA and year fixed effects, as well as MSA-year unemployment rates and male wages. Data sources are: Vital Statistics (births, population), Census and Federal Housing Finance Agency (house prices), Current Population Survey (wages, fraction college), and Bureau of Labor Statistics (unemployment rates). All specifications are weighted by the total number of births in the cell. Standard errors adjusted for clustering at the MSA level are in parentheses. * $p<.1$, ** $p<.05$, *** $p<.01$.

from an MSA-group with an ownership rate of 0 to a cell with an ownership rate of 1, there would be a relative increase of 2.5% in fertility rates. More usefully, if house prices increase by \$10,000, comparing MSA-groups with ownership rates of 0.25 to those with ownership rates of 0.75, we would see a relative increase of 1.25% in fertility rates.

4.2. Instrumental variable specifications

The main threat to assigning a causal interpretation to the estimated β_1 and β_2 is the possibility of reverse causality or some correlated unobservable to house prices that affect fertility rates. If it were simply the case that in MSAs where people demanded more children house prices were driven up in equilibrium, ceteris paribus, then both β_1 and β_2 would be estimated to be positive. For the finding of separating effects to be explained by the alternative reverse causality story, it must be the case that fertility-related demand pressures occur disproportionately in areas with relatively higher rates of home ownership (as measured in a pre-period baseline year). This confounding story is one of fertility-preference demand driven price changes.

In order to address the possibility that reverse causality or correlated unobservables are biasing our estimates, we make use of the Saiz (2010) measure of housing supply elasticity. The measure is based on non-linear combinations of both the Saiz (2008) geographic limitations measure and the Wharton Residential Urban Land Regulation Index created by Gyourko et al. (2008). We propose that concerns about fertility-preference demand driven price changes are less likely to be a concern in places with lower housing supply constraints, or higher housing market supply elasticities. We thus estimate our regression models separately for MSAs with higher and lower levels of supply elasticity, as captured by the Saiz (2010) measure. If the estimated relationship is maintained in less supply constrained places, that bolsters our confidence that our estimated effect is not driven by homeowners with infants (or fertility intentions) bidding up the prices of inelastically supplied houses.

Table 4 reports the results using the baseline specification displayed in column 2 of Table 3, which include both MSA and year fixed effects. We choose this to be our baseline specification because it is the one we will estimate with the IV specifications, as described below. Column 1 reports the results for the sample of MSAs with supply elasticities below the median and column 2 above the median. In fact, moving from column 1 to 2, the estimated positive coefficient on $HousePrice_{mt-1} * OwnRate_{mg}$ increases in magnitude, as does the estimated conditional negative main effect of $HousePrice_{mt-1}$. This is opposite of what would be expected under the reverse causality scenario.

Next, we more formally incorporate the supply elasticity measure by employing an instrumental variables strategy similar to that employed by Chetty and Szeidl (2012); we instrument for MSA-level house prices with the interaction of a baseline supply elasticity measure with the national trend in housing prices. The intuition here is that aggregate demand shocks that affect the national housing market are expected to exert a relatively larger influence on local housing prices in MSAs which are more supply constrained. The identification assumption is that the interaction between baseline MSA housing supply elasticity and national house price trends would not have been systematically correlated with trends in fertility rates in the absence of MSA house price changes.

In order to implement this strategy we interact the measure of supply elasticity with the national version of FHFA housing price index. Since we have two potentially endogenous variables – the level measure of house prices and the interaction term with MSA-group-level ownership rates – we use the triple interaction of the MSA supply elasticity with the national house price index and MSA-group-level ownership rate as a second instrumental variable. Since nationally, house prices grew almost linearly in this time period, we do not include MSA-specific time trends as our baseline IV specification.[15] We also do not include the conditional main effect of housing supply elasticity because it is measured at the baseline and does not vary across groups, so it is absorbed by the MSA fixed effects. Table 4 presents the first-stage estimation results for the interaction term $Houseprice_{mt-1} * OwnRate_{mg}$ (column 3) and the level term $Houseprice_{mt-1}$ (column 4). As expected, increasing supply elasticity is associated with reductions in MSA-level house price growth as national house prices increase.

Table 4, column 5 presents the IV results. The estimated effects of interest maintain their signs of direction, but increase in magnitude. The point estimate of β_1 of 0.0723 and a point estimate on β_2 of -0.0239, are both statistically significant at the one percent level. These estimates suggest that if house prices increase by \$10,000, as we move from an MSA-group with an ownership rate of 0 to a cell with an ownership rate of 1, there would be a relative increase of 7.2% in fertility rates. More usefully, if house prices increase by \$10,000, comparing MSA-groups with ownership rates of 0.25 to those with ownership rates of 0.75, we would see a relative increase of 3.6% in fertility rates. These estimates are larger in magnitude than the OLS results and a Hausman specification test can reject the consistency of the OLS estimate at the 1% level.[16] However, the net effect at the mean is almost identical between the OLS and IV specification: both specifications indicate that at the mean U.S. home ownership rate, the net effect of a \$10,000 increase in house prices is a 0.8% increase in fertility rates. We put these numbers into context below with the use of simulation exercises.

4.3. Robustness checks

In this section we implement various robustness checks on the model specification and sample construction. We begin by considering how our estimates change if we replace the house price in the year of conception with alternative measures of house prices. We do not have a strong reason to believe that house prices in the year of conception is the most relevant measure, as opposed to, say, house prices averaged over the three years prior. It may be the case that couple's fertility decisions are based on a longer time horizon or on longer terms averages. Table 5 reports the results of estimating alternative models of this sort, using both the OLS and IV strategies. Specifications in columns 1–4 use house prices in the years 1, 2, 3, and 4, respectively, prior to conception. Specifications in columns 5, 6, and 7 use the 3-year moving average of house prices over the two, three, and four years, respectively, prior to conception. In all of these seven alternative models, the familiar pattern emerges of a positive coefficient on the interaction between $HousePrice_{mt-1}$ and $OwnRate_{mg}$ and a negative coefficient on $HousePrice_{mt-1}$, for both the OLS and IV results. The point estimates of β_1 ranging from 0.0468 to 0.0718 for the OLS and 0.0723 to 0.0995 for the IV results.

Next, we estimate various alternative specifications to Eq. (1) above, providing some robustness checks on the main MSA-group-level analysis. Table 6 reports these results. Column 1 reproduces the main OLS and IV results from Tables 3 and 4 for the sake of comparison. In column 2 we replace male wages with separate measures of male and female wages. In column 3 we replace the wage distribution measures with the mean wage. In column 4 we replace the wage distribution measures with a measure of income per capita collected from the Bureau of Economic Analysis (BEA) regional economic accounts. To create this variable at the MSA-year level, we employ our crosswalk procedure described in the Appendix A. In each case, the coefficients are virtually unchanged.

In column 5 we consider that owners and non-owners might be differentially affected by general economic conditions in a way that is not captured by simply including a measure of wages. If this were the case, the coefficient on $HousePrice_{mt-1} * OwnRate_{mg}$ might capture this difference, leading to a biased estimate of the causal effect of interest. To do this, we interact the home-ownership rate with the wage measures. Column 5 displays the results of this exercise. The coefficient on $75thWage_{mt-1} * OwnRate_{mg}$ is positive and statistically significant, indicating owner's fertility decisions are positively affected by increases in male wages at the top of the distribution. However, the coefficients on $HousePrice_{mt-1} * OwnRate_{mg}$ and $HousePrices_{mt-1}$ remain unaffected.[17]

4.4. Different demographic groups

In Table 7, we report the results of estimating Eq. (1) for various demographic subgroups and for first and higher order births using both the OLS and IV strategies. Column 2 reports the results for Non-Hispanic Whites, column 3 reports the results for Non-Hispanic Blacks and column 4 reports the results for Hispanic Whites. The point estimate on the interaction term $HousePrice_{mt-1} * OwnRate_{mg}$ is always positive while the coefficient on $HousePrice_{mt-1}$ is negative, implying a net positive effect of house price increases among home owners and a negative effect among non-owners across all groups.

We next consider whether the effects of house prices on current period births are driven by first births, second births, or higher parity births. It is not clear a priori which would be more price or income elastic. On the one hand, the optimal timing of first births might be less constrained, since mothers tend to be younger and might consider that a deliberate delay will be less consequential, as they have more childbearing years ahead of them. Also, if couples have specific ideas about optimal spacing, they might be more flexible about the timing of their first birth. On the other hand, subsequent births might be more "marginal" and thus might exhibit a great degree of elasticity with respect to price or a wealth shock. An additional motivation for this analysis is that an effect on higher order births might be indicative of a change in completed fertility.

Table 7 columns 5–7 report the results. For both first, second, and higher parity births, the estimated coefficient on the interaction between $HousePrice_{mt-1}$ and ownership rate is positive and statistically significant, with similar magnitudes: 0.0538, 0.0474 and 0.0434 in the OLS specification and 0.0947, 0.0754 and 0.0574 in the IV specification, respectively. The point estimate for the coefficient on $HousePrice_{mt-1}$ is negative and statistically significant for first, second, and higher-order births. The finding of an effect on both first and higher-order births is potentially informative about the nature of the effects we are estimating. Increases in first births might reasonably be interpreted as a change in timing, while changes in higher order births might reasonably be interpreted as an increase in the total number of children, particularly for third and higher parity births. These interpretations are merely speculative, and warrant further investigation.

Given that a previous literature exists on the relationship between unemployment rates and contemporaneous fertility rates, it is interesting to consider the estimated coefficients on the unemployment rate. Our

[15] If we include MSA time trends, the results are very similar but the first stage F Statistics fall below conventional levels. The coefficients on β_1 and β_2 are 0.07148 and -0.01439 (with standard errors of 0.0098 and 0.0039), respectively. However, the first stage F Statistic is the equation for $Houseprice_{mt-1}$ is reduced to 8.71, which is below the conventional rule of thumb of 10, as well as the Stock–Yogo critical values (Stock and Yogo, 2005).

[16] Since standard errors are adjusted for clustering in both the OLS and 2SLS specifications, OLS is not fully efficient and we compute the Hausman test statistic with bootstrapped variance estimates. The variance was calculated using 500 bootstrap replications.

[17] We performed three additional noteworthy robustness checks, but do not report them for sake of space: (1) add a control variable for average rental prices in the MSA-year; (2) consider an alternative sample of MSAs that did not change boundaries between 1990 and 2009; and (3) use only variation across MSA/age category cells to define average ownership rates. None of these changes alters the estimated coefficients of interest in a meaningful way.

Table 5
Alternative house price measures.

Dep. Var. $log(FertRate)_{mgt}$	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	$Price_{mt-1}$	$Price_{mt-2}$	$Price_{mt-3}$	$Price_{mt-4}$	$AvgPrice_{mt-2}$	$AvgPrice_{mt-3}$	$AvgPrice_{mt-4}$
OLS							
$HousePrice_{mt-1} * OwnRate_{mg}$	0.0468***	0.0534***	0.0630***	0.0718***	0.0501***	0.0541***	0.0581***
	(0.0045)	(0.0052)	(0.0063)	(0.0077)	(0.0048)	(0.0052)	(0.0057)
$HousePrice_{mt-1}$	−0.0128***	−0.0145***	−0.0172***	−0.0199***	−0.0136***	−0.0146***	−0.0157***
	(0.0012)	(0.0013)	(0.0017)	(0.0023)	(0.0012)	(0.0014)	(0.0015)
IV							
$HousePrice_{mt-1} * OwnRate_{mg}$	0.0723***	0.0823***	0.0919***	0.0995***	0.0770***	0.0814***	0.0853***
	(0.0100)	(0.0111)	(0.0120)	(0.0127)	(0.0105)	(0.0109)	(0.0113)
$HousePrice_{mt-1}$	−0.0239***	−0.0274***	−0.0313***	−0.0353***	−0.0255***	−0.0272***	−0.0288***
	(0.0033)	(0.0036)	(0.0040)	(0.0045)	(0.0034)	(0.0036)	(0.0038)
$OwnRate_{mg}$	−0.171	−0.301	−0.423	−0.509*	−0.232	−0.288	−0.336
	(0.291)	(0.282)	(0.275)	(0.271)	(0.287)	(0.283)	(0.280)
$UnempRate_{mt-1}$	−0.0137**	−0.0137**	−0.0120**	−0.00893*	−0.0138**	−0.0134**	−0.0126**
	(0.00563)	(0.00573)	(0.00547)	(0.00494)	(0.00567)	(0.00559)	(0.00542)
$25thWage_{mt-1}$	0.00226	0.00222	0.00211	0.00202	0.00224	0.00220	0.00216
	(0.00166)	(0.00164)	(0.00155)	(0.00146)	(0.00165)	(0.00162)	(0.00158)
$50thWage_{mt-1}$	0.00153	0.00145	0.00134	0.00137	0.00149	0.00144	0.00142
	(0.00115)	(0.00115)	(0.00109)	(0.00103)	(0.00115)	(0.00112)	(0.00109)
$75thWage_{mt-1}$	0.00120*	0.00121*	0.00120*	0.00113**	0.00120*	0.00120*	0.00118*
	(0.000704)	(0.000701)	(0.000649)	(0.000561)	(0.000702)	(0.000685)	(0.000653)
$FracColl_{mgt-1}$	−0.492***	−0.505***	−0.502***	−0.491***	−0.498***	−0.499***	−0.497***
	(0.102)	(0.102)	(0.0991)	(0.0951)	(0.102)	(0.101)	(0.0997)
Group fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
MSA fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
F Stat 1	29.60	27.83	25.94	23.78	28.87	28.24	27.67
F Stat 2	18.32	18.69	18.84	18.23	18.61	18.91	19.27
Number of MSAs	154	154	154	154	154	154	154
N	9240	9240	9240	9240	9240	9240	9240

Notes: Fertility rates are total births over the total female population in each MSA, year of conception, age category and race/ethnicity cell for women aged 20–44. Mean home ownership rates are calculated in 1990 Census by year, MSA, age category, and race/ethnicity. Fraction of cell that is a college graduate is matched by MSA, year, age category and race. House prices (10,000 s) are matched by MSA and year of conception (or years prior to conception where noted). Average house price refers to the average home price from the year indicated up to the year of conception. The instrumental variable is the interaction between Saiz (2010) supply elasticity measure and the national version of the FHFA house price index. First stage1 refers to the first stage regression where the dependent variable is $HousePrice_{mt-1} * OwnRate_{mg}$ and first stage 2 refers to the first stage regression where the dependent variable is $HousePrice_{mt-1}$. F Stat 1 refers to the first stage F statistic where the dependent variable is $HousePrice_{mt-1} * OwnRate_{mg}$, and F Stat 2 for $HousePrice_{mt-1}$. Data sources are: Vital Statistics (births, population), Census and Federal Housing Finance Agency (house prices), Current Population Survey (wages, fraction college), and Bureau of Labor Statistics (unemployment rates). All specifications are weighted by the total number of births in the cell. Standard errors adjusted for clustering at the MSA level are in parentheses. * $p < .1$, ** $p < .05$, *** $p < .01$.

regression models yield statistically significant negative estimates of the relationship between unemployment rates and fertility rates among Whites, but not among Blacks or Hispanics. When house prices are not included in the model (not shown in the table), the estimated relationship is largely unchanged for Whites (a statistically insignificant −0.0036), but it becomes positive and statistically significant for Blacks and Hispanics. It is also interesting to note that in terms of separate effects by birth parity, the unemployment rate is negatively related to first and second births, but not discernibly related to higher-order births. This would be consistent with the unemployment rate having an effect on the timing of childbearing initiation, but potentially not with completed fertility. To the extent that this interpretation is warranted, this is an interesting contrast to the potentially more permanent effect of house prices. Again, we think these considerations deserve further examination, although it is outside the scope of this paper.

4.5. Individual level estimation

The empirical results presented above suggest that an increase in MSA-level house prices exert a negative price effect on births among non-owners and a net positive effect on births among owners, all else equal. These estimates are generated by an aggregated cell-level analysis, but the underlying conceptual framework is at the individual level. We thus turn to individual-level Current Population Survey (CPS) data to check that the story told by aggregate level data is confirmed with individual level data. We map the older MSA designations provided in the CPS (as in the Census) to the 2009 MSA designations provided in the FHFA house price data using the crosswalk procedure described in the Appendix A. In the CPS we do not see the full population of births, as we do with an analysis of Vital Statistics birth data. However, as a supplementary data source, the CPS offers the distinct advantage of directly identifying home-owners.

In this individual level analysis, we define own_i as an indicator for whether the individual in the CPS is the household head or head's spouse and the household is reported to own their home. In the aggregate analysis above, ownership was defined at the group level in the baseline year of 1990. Caution should thus be exercised in assigning a causal interpretation to the $HousePrice_{mt-1} * Own_i$ interaction term in this specification, since individuals who intend to have a baby this year might decide to buy a house in anticipation of that event. This is another reason we consider this analysis supplementary to the main analysis above.

We define the dependent variable $Pr(Birth)_i$ to equal one if there is a child under the age of one in the household. We include as controls a set of indicator variables for the mother's age (in categories), race/ethnicity and level of education. All the other variables are defined at the MSA level as defined in Eq. (1) above. Explanatory variables, including the house price index, are matched to observations by the year prior to the survey year in order to capture the effect of conditions in the year of the baby's conception. (We do not have perfect birth-date or gestation information, as we do in the Vital Statistics Natality Files, and so here we use year minus one as an approximation.) Table 8 reports the results estimated using a linear probability model. In the pooled sample regression reported in column 1, we see the familiar pattern of point estimates — a negative point estimate on $HousePrice_{mt-1}$ and a positive point estimate on the interaction of $HousePrice_{mt-1} * Own_i$ (significant at the 1% level).

Table 6
Alternative controls.

Dep. Var. log(FertRate)$_{mgt}$	(1)	(2)	(3)	(4)	(5)
OLS					
$HousePrice_{mt-1} * OwnRate_{mg}$	0.0468***	0.0468***	0.0468***	0.0468***	0.0389***
	(0.00448)	(0.00447)	(0.00447)	(0.00446)	(0.00497)
$HousePrice_{mt-1}$	−0.0128***	−0.0128***	−0.0127***	−0.0128***	−0.0101***
	(0.00115)	(0.00115)	(0.00116)	(0.00119)	(0.00123)
IV					
$HousePrice_{mt-1} * OwnRate_{mg}$	0.0723***	0.0723***	0.0723***	0.0722***	0.0739***
	(0.00997)	(0.00997)	(0.00996)	(0.00995)	(0.0128)
$HousePrice_{mt-1}$	−0.0239***	−0.0239***	−0.0239***	−0.0240***	−0.0244***
	(0.00328)	(0.00325)	(0.00328)	(0.00334)	(0.00412)
$OwnRate_{mg}$	−0.171	−0.171	−0.171	−0.171	−0.0915
	(0.291)	(0.292)	(0.291)	(0.291)	(0.349)
$UnempRate_{mt-1}$	−0.0137**	−0.0139**	−0.0138**	−0.0122**	−0.0138**
	(0.00563)	(0.00568)	(0.00573)	(0.00530)	(0.00570)
$FracColl_{mgt-1}$	−0.492***	−0.491***	−0.490***	−0.486***	−0.493***
	(0.102)	(0.102)	(0.101)	(0.101)	(0.0977)
$25thWage_{mt-1}$	0.00226				−0.00341
	(0.00166)				(0.00769)
$50thWage_{mt-1}$	0.00153				0.00752
	(0.00115)				(0.00550)
$75thWage_{mt-1}$	0.00120*				0.000790
	(0.000704)				(0.00501)
$25thWageAll_{mt-1}$		0.00468**			
		(0.00231)			
$50thWageAll_{mt-1}$		−0.000306			
		(0.00185)			
$75thWage_{mt-1}$		0.00262**			
		(0.00121)			
$MeanWage_{mt-1}$			0.00329***		
			(0.00120)		
$IncomePC_{mt-1}$				0.00603*	
				(0.00328)	
$25thWage_{mt-1} * OwnRate_{mg}$					0.0147
					(0.0189)
$50thWage_{mt-1} * OwnRate_{mg}$					−0.0156
					(0.0148)
$75thWage_{mt-1} * OwnRate_{mg}$					0.00106
					(0.0124)
Group fixed effects	Yes	Yes	Yes	Yes	Yes
MSA fixed effects	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes
F Stat 1	29.60	29.76	29.81	28.09	28.27
F Stat 2	18.32	18.26	18.44	16.09	18.25
No. of MSAs	154	154	154	154	154
N	9240	9240	9240	9240	9240

Notes: Fertility rates are total births over the total female population in each MSA, year of conception, age category and race/ethnicity cell for women aged 20–44. Mean home ownership rates are calculated in 1990 Census by year, MSA, education category, age category, and race. Fraction of cell that is a college graduate is matched by MSA, year, age category and race. House prices (10,000 s), Income per capita, male wages, and all wages are matched by MSA and year of conception. The instrumental variable is the interaction between Saiz (2010) supply elasticity measure and the national version of the FHFA house price index. First stage 1 refers to the first stage regression where the dependent variable is $HousePrice_{mt-1} * OwnRate_{mg}$ and first stage 2 refers to the first stage regression where the dependent variable is $HousePrice_{mt-1}$. Data sources are: Vital Statistics (births, population), Census and Federal Housing Finance Agency (house prices), Current Population Survey (wages, fraction college), and Bureau of Labor Statistics (unemployment rates). All specifications are weighted by the total number of births in the cell. Standard errors adjusted for clustering at the MSA level are in parentheses. * $p < .1$, ** $p < .05$, *** $p < .01$.

Columns 2–3 report the results including the additional MSA and year fixed effects and the time-varying MSA-level controls, unemployment rates and wages. Column 4 reports the results from the IV specification described above. Although the IV results are not precisely estimated, the magnitudes of the coefficients of primary interest are similar to the OLS estimates. Overall, this set of individual-level results gives us confidence that our interpretation of the results from the aggregate level analyses is appropriate. In particular, we see that the positive effect is being driven by individuals that are self-reported to be home owners.

4.6. Housing bust periods (1990–1996 and 2007–2009)

Our analysis has thus far been limited to a period of history characterized by rising house prices. It is interesting to consider explicitly the relationship between housing price decreases and birth rates. There might be asymmetric effects, whereby an increase in housing wealth might lead people move up their period of childbearing to a greater extent than a decrease in housing wealth will lead people to delay. One possible reason for such an asymmetry is that there is a biological timing constraint that individuals are reluctant to push against. It becomes an empirical question as to whether there are differential responses to house price rises and declines. To consider this explicitly, we use data from two periods of house price decline: 1990–1996 and 2007–2010. Fig. 1 shows these two periods: between 1990 and 1996 prices declined gradually and between 2007 and 2010 there is a dramatic decline in prices. Unfortunately Vital Statistics birth data is not yet available for conception years past 2006, so we can only look at the 1990–1996 housing bust period using the approach of the main analysis. We therefore turn to individual level data sources for 2007–2010 period.

We begin by examining the 1990–1996 bust period using the approach used in the main analysis. Table 9, columns 1 and 2 display the results using OLS and the IV strategy, respectively. The pattern on the coefficients remains similar to the 1997–2006 period — a positive coefficient on $HousePrice_{mt-1} * Own_{mg}$ and a negative coefficient

Table 7
Different groups.

Dep. Var. log(FertRate)$_{mgt}$	(1) All	(2) White	(3) Black	(4) Hispanic	(5) First births	(6) Second births	(7) Higher births
OLS							
$HousePrice_{mt-1} * OwnRate_{mg}$	0.0468***	0.0733***	0.0788***	0.0220***	0.0538***	0.0474***	0.0434***
	(0.00448)	(0.00762)	(0.00922)	(0.00302)	(0.00522)	(0.00409)	(0.00551)
$HousePrice_{mt-1}$	−0.0128***	−0.0306***	−0.0142***	−0.00690***	−0.0122***	−0.0118***	−0.0152***
	(0.00115)	(0.00286)	(0.00141)	(0.00196)	(0.00121)	(0.00117)	(0.00169)
IV							
$HousePrice_{mt-1} * OwnRate_{mg}$	0.0723***	0.106***	0.0926*	0.0305***	0.0947***	0.0754***	0.0574***
	(0.00997)	(0.0133)	(0.0483)	(0.00999)	(0.0129)	(0.0114)	(0.00988)
$HousePrice_{mt-1}$	−0.0239***	−0.0485***	−0.0201***	−0.0116***	−0.0260***	−0.0230***	−0.0256***
	(0.00328)	(0.00598)	(0.00694)	(0.00339)	(0.00386)	(0.00372)	(0.00382)
$OwnRate_{mg}$	−0.171	1.746***	−2.332***	−0.746***	−0.149	−0.0477	−0.423
	(0.291)	(0.613)	(0.486)	(0.141)	(0.311)	(0.415)	(0.299)
$FracColl_{mgt-1}$	−0.492***	−0.248**	−0.120**	−0.0279	−0.462***	−0.533***	−0.625***
	(0.102)	(0.120)	(0.0562)	(0.0496)	(0.140)	(0.128)	(0.120)
$UnempRate_{mt-1}$	−0.0137**	−0.0158**	0.000361	0.00269	−0.0237***	−0.0111**	−0.00947
	(0.00563)	(0.00703)	(0.00810)	(0.00496)	(0.00631)	(0.00513)	(0.00757)
$25thWage_{mt-1}$	0.00226	0.00131	0.00350*	0.000918	0.00294	0.00185	0.00254
	(0.00166)	(0.00157)	(0.00186)	(0.00302)	(0.00204)	(0.00155)	(0.00223)
$50thWage_{mt-1}$	0.00153	0.00118	0.000151	0.00185	0.00177	0.00126	0.00168
	(0.00115)	(0.00131)	(0.00145)	(0.00297)	(0.00122)	(0.00112)	(0.00185)
$75thWage_{mt-1}$	0.00120*	0.00111	0.00114**	0.000876	0.000955	0.00109**	0.00164
	(0.000704)	(0.000944)	(0.000499)	(0.00158)	(0.000734)	(0.000553)	(0.00121)
Group fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
MSA fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
F Stat 1	29.60	24.57	15.81	22.57	29.54	30.59	27.60
F Stat 2	18.32	18.58	17.19	20.71 19.26	18.86	16.32	
N	9240	3080	3080	3080	9234	9240	9240

Notes: Fertility rates are total births over the total female population in each MSA, year of conception, age category and race/ethnicity cell for women aged 20–44. Mean home ownership rates are calculated in 1990 Census by year, MSA, age category, and race/ethnicity. House prices (10,000 s), unemployment rates, and male wages are matched by MSA and year of conception. All specifications include a control for the fraction of women who are college educated in the MSA-group-year. The instrumental variable is the interaction between Saiz (2010) supply elasticity measure and the national version of the FHFA house price index. First stage 1 refers to the first stage regression where the dependent variable is $HousePrice_{mt-1} * OwnRate_{mg}$ and first stage 2 refers to the first stage regression where the dependent variable is $HousePrice_{mt-1}$. Data sources are: Vital Statistics (births and population), Census and Federal Housing Finance Agency (house prices), Current Population Survey (wages), and Bureau of Labor Statistics (unemployment rates). All specifications are weighted by the total number of births in the cell. Standard errors adjusted for clustering at the MSA level are in parentheses. * $p<.1$, ** $p<.05$, *** $p<.01$.

$HousePrice_{mt-1}$. In columns 3 and 4 we report results from an individual level analysis from the CPS for this time period. Again, results are very similar to those found in the 1997–2006 period: a positive and statistically coefficient on $HousePrice_{mt-1} * Own_{mg}$ and a negative, but not statistically significant effect on $HousePrice_{mt-1}$. As in the analysis for 1997–2006 period, the IV estimates are similar but less precise.

Next, we move on to the 2007–2010 housing bust period, which is characterized by a steep decline in prices. First, we repeat the individual-level CPS analysis for this period. Table 9, columns 5 and 6 display results. The pattern on the coefficients is extremely similar to both the earlier bust period (1990–1996) and the housing boom period (1997–2006). Since Vital Statistics birth data is not available for this period, we supplement the analysis by examining data from the American Communities Survey (ACS), conducted annually by the U.S. Census Bureau, beginning in 2000. We obtained this data from IPUMS. The data is available with the equivalent of MSA identifiers starting in 2005.[18] We construct the indicator variables "Pr(Birth)" and "own home" in the same manner as described above for the CPS data. Again, the coefficient on $HousePrice_{mt-1} * Own_i$ is positive and statistically significant at the 1% level. The coefficient on $HousePrice_{mt-1}$ is positive, although it is not statistically significant.

These findings give us some confidence that it is appropriate to use our preferred aggregate results above – generated from data for the years 1997–2006, a period characterized by house price increases — to make out-of-sample predictions to more recent years, characterized by house price declines. Between 2006 and 2010, housing prices fell $63,000 among the MSAs in our sample. At the mean rate of home ownership, our estimates imply that this would lead to a 7.5% decline in birth rates. We can also simulate the effect of the rise in unemployment rates over the period.[19] Between 2006 and 2010, unemployment rates rose 5.14 percentage points. Holding housing prices fixed, our estimates imply that this corresponds to a 6.8% decline in births. According to the National Center for Health Statistics, the national fertility rate dropped from 69.3 in 2007 to 63.8 in 2011, a 7.9% decline.

4.7. Interpreting the magnitudes of the estimated effects

Our analysis of Vital Statistics birth data coupled with MSA-level house prices shows that an increase in MSA-level house prices, all else held constant, is associated with fewer births among non-owners and a net increase in births among owners. We interpret this pair of findings as indicative of a negative price effect among non-owners and a dominant housing wealth/equity effect among owners. These patterns hold for Whites, Blacks, and Hispanics and appear for both first and higher parity births.

In order to facilitate an understanding of whether these results are economically large or small, we conduct a simple simulation exercise. Fig. 3 presents the predicted effect of a $10,000 increase in house prices on births for each race/ethnic group as well as first and higher parity

[18] The ACS identifies PUMAs (Public Use Microdata Areas), which IPUMS has matched to MSAs. We then use the crosswalk procedure described in the data appendix to match to the housing data (Ruggles et al., 2010). PUMAs are also identified in 2003, but we do not use this data.

[19] Both the average fall in home prices and the average increase in unemployment rates are population weighted average changes for the 154 MSAs in our sample between 2006 and 2010.

Table 8
Individual level analysis using CPS: 1997–2006.

Dep. Var. $Pr(Birth)_i$	(1)	(2)	(3)	(4)
	No MSA/year FE	MSA/year FE	MSA/year FE	IV
$HousePrice_{mt-1} * Own_i$	0.000631***	0.000569***	0.000567***	0.000427
	(0.000197)	(0.000197)	(0.000197)	(0.000409)
$HousePrice_{mt-1}$	−0.000134	−0.000307**	−0.000331**	−0.0000331
	(0.0000985)	(0.000131)	(0.000138)	(0.000337)
Own_i	0.0367***	0.0377***	0.0378***	0.0403***
	(0.00385)	(0.00393)	(0.00393)	(0.00782)
$UnempRate_{mt-1}$			0.000156	0.000541
			(0.000650)	(0.000735)
$25thWage_{mt-1}$			−0.000382	−0.000350
			(0.000639)	(0.000637)
$50thWage_{mt-1}$			0.000361	0.000338
			(0.000551)	(0.000551)
$75thWage_{mt-1}$			0.000229	0.000197
			(0.000274)	(0.000280)
Demographics	Yes	Yes	Yes	Yes
MSA fixed effects	No	Yes	Yes	Yes
Year fixed effects	No	Yes	Yes	Yes
IV	No	No	No	Yes
Mean had baby	0.063	0.063	0.063	0.063
Mean own	0.498	0.498	0.498	0.498
F Stat 1				25.64
F Stat 2				17.08
N	192788	192788	192788	192788

Notes: Sample is women aged 20–44 in March Current Population Survey 1998–2007. Dependent variable is an indicator for having a child under one. House prices (10,000 s), unemployment rates, and male wages are matched by MSA and year. Ownership is the household's home ownership status, which is assigned as a 1 when the household owns a home and the respondent is the household head or spouse of the household head. Demographic controls include fixed effects for education, year, age category, and race/ethnicity. The instrumental variable is the interaction between Saiz (2010) supply elasticity measure and the national version of the FHFA house price index. First stage 1 refers to the first stage regression where the dependent variable is $HousePrice_{mt-1} * OwnRate_{mg}$ and first stage 2 refers to the first stage regression where the dependent variable is $HousePrice_{mt-1}$. Data sources are: Census and Federal Housing Finance Agency (house prices), Current Population Survey (individual level data, wages), and Bureau of Labor Statistics (unemployment rates). Standard errors adjusted for clustering at the MSA level are in parentheses. * $p<.1$, ** $p<.05$, *** $p<.01$.

births.[20] The x-axis represents group home ownership rates and the y-axis represents the net predicted percentage change in births from of a $10,000 increase in house prices, conditional on each level of home ownership. The prediction is indicated by the solid line and a 95% confidence interval is indicated by the dashed lines.[21] The predictions are calculated based on IV point estimates displayed in Table 7, which include all of the main demographic group and MSA-level control variables, and MSA and year fixed effects.

In all cases, the exercise suggests a positive, linear relationship between home ownership rates and the change in births due to a $10,000 increase in house prices. The net effect for all demographic groups implies that as the ownership rate increases from 30% to 40%, the net effect become positive. This implies that in MSAs with sizable rates of home ownership, the positive home equity effect among owners is large enough to outweigh the negative price effect, leading to increases in MSA-level birth rates.

We also consider what changes in home prices imply for group specific fertility rates, since there is heterogeneity in the magnitude of the price and home equity effects as well as in rates of home ownership. Overall in our data the population weighted mean home ownership rate is 44%. At this rate, the net effect of a $10,000 increase in prices is a 0.8% increase in births. Among Whites, the mean home ownership rate is 53%, which is associated with a 0.7% increase in births for that group. Among Blacks, the mean home ownership rate is 24%, which is associated with a net increase of 0.2% in births. And among Hispanics, the mean home ownership rate is 29%, which is associated with a net decrease in births of 0.3%. This indicates that although Hispanic home ownership rates are higher than Black home ownership rates, the net effect is smaller for Hispanics because the estimated home equity effect is smaller for that group.

Finally, it is useful to consider an out-of-sample prediction assuming extreme values of ownership rates, to have an estimate of the effect among owners and non-owners. Assuming a 100% ownership rate, the net impact of a $10,000 house price increase is a 5% increase overall. Separately by race/ethnicity, our simulations suggest a 5.9% increase for Whites, a 7.9% increase for Blacks, and a 1.9% increase for Hispanics. These figures imply that among owners, the increase in house prices during the recent housing boom led to a sizable impact on the likelihood of giving birth in a given year.[22]

[20] Since the effects are similar for second and third or higher parity births we combine the two for these simulation exercises.

[21] We predict the percentage change in fertility rates from a $10,000 increase in mean housing prices for ownership rate o: $(FertRate|HousePrice = h + 10k, OwnRate = o) - (FertRate|HousePrice = h, OwnRate = o)/(FertRate|HousePrice = h, OwnRate = o)$. For each group, we calculate the standard error of the prediction at the mean of the independent variables using 100 bootstrap replications and apply that standard error to calculate the confidence interval at each level of o. The solid line represents the predicted effect and the dashed line represents a 95% confidence interval, both of which were smoothed using a locally weighted linear regression.

[22] These results are comparable to those found in a contemporaneous working paper by Lovenheim and Mumford (2011), which investigates the relationship between changes in home value and current period fertility using individual-level data from the Panel Study of Income Dynamics (PSID) from 1990 to 2007. The authors estimate linear probability models of the probability that a woman gives birth in a given year as a function of two and four year changes in the reported market value of her home. The authors find that a $10,000 increase in an individual's real housing wealth is associated with a 0.07 percentage point (1.3% at the mean) increase in the probability of having a child. It is also useful to compare our estimates to those found by, in their analysis of the effect of earnings on current period birth rates. Those authors use the experience of the coal boom and bust during the 1970s and 1980s in the Appalachian region of the U.S. to examine the effect of an exogenous increase in male's earnings (because females almost never work in the coal industry) on fertility rates. They estimate changes in county-level birth rates as a function of differences in lagged county level log earnings, conditional on state and year fixed effects. They find that a ten percent increase in county-level earnings is associated with a one percent increase in the subsequent year's birth rates. For earning increases coming specifically from the coal boom, they estimate that a ten percent increase in coal-related earnings is associated with a seven percent increase in the subsequent year's birth rates query. To put our findings in comparable terms, recall that we simulated that a $10,000 increase in house prices is associated with a 0.8% net increase in birth rates. Consider that a $10,000 increase in house prices is about a five percent increase off the 2006 median house price in the U.S. of roughly $260,000. So our estimates would suggest that a 10% increase in house prices would be associated with a 1.6% net increase in birth rates.

Table 9
Real estate bust periods.

Dep. Var., Source	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	OLS	IV	OLS	IV	OLS	IV	OLS	IV
	log(fert rate), aggregate-90-96		Pr(birth), CPS-90-96		Pr(birth), CPS-07-09		Pr(birth,) ACS-07-09	
$HousePrice_{mt-1} * Own$	0.0604***	0.0722***	0.000707*	0.000741	0.000703***	0.000978***	0.000624***	0.000726***
	(0.00765)	(0.00962)	(0.000402)	(0.000587)	(0.000224)	(0.000353)	(0.000119)	(0.000188)
$HousePrice_{mt-1}$	−0.0117**	−0.0206***	−0.000560	−0.000379	−0.000689	−0.00110	−0.0000414	0.000326
	(0.00526)	(0.00732)	(0.000501)	(0.00199)	(0.000480)	(0.00122)	(0.000241)	(0.000342)
Own	0.103	0.0170	0.0362***	0.0357***	0.0318***	0.0258***	0.0407***	0.0385***
	(0.365)	(0.343)	(0.00511)	(0.00770)	(0.00496)	(0.00711)	(0.00206)	(0.00318)
$UnempRate_{mt-1}$	−0.00761**	−0.0127*	−0.000774	−0.000649	−0.000322	−0.000763	0.000617	0.00119**
	(0.00306)	(0.00698)	(0.000971)	(0.00155)	(0.00145)	(0.00220)	(0.000482)	(0.000590)
$25thWage_{mt-1}$	−0.000971	−0.00130	−0.000730	−0.000725	0.000577	0.000571	0.0000121	0.00000597
	(0.00125)	(0.00151)	(0.000753)	(0.000752)	(0.00128)	(0.00128)	(0.000356)	(0.000348)
$50thWage_{mt-1}$	0.00203*	0.00269*	0.000514	0.000502	−0.000342	−0.000383	−0.000330	−0.000261
	(0.00121)	(0.00138)	(0.000726)	(0.000733)	(0.00118)	(0.00117)	(0.000351)	(0.000352)
$75thWage_{mt-1}$	0.000598	0.000663	0.000464	0.000464	−0.000485	−0.000458	0.000216	0.000197
	(0.000796)	(0.000745)	(0.000487)	(0.000485)	(0.000602)	(0.000605)	(0.000176)	(0.000181)
Demographic fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
MSA fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Mean fert rate	0.069	0.069	0.064	0.064	0.066	0.066	0.058	0.058
Mean own	0.45	0.45	0.43	0.43	0.46	0.46	0.45	0.45
F stat 1		22.55		15.88		36.93		45.43
F stat 2		8.48		3.94		13.88		27.53
N	6348	6348	126799	126799	71317	71317	801798	801798

Notes: In columns (1)–(2) the sample is all births to women aged 20–44 for the bust period of 1990–1996 according to the specification in Table 3, column (6) and the dependent variable is the fertility rate. In columns (3)–(4), the sample is women aged 20–44 in the March Current Population Survey for the bust period 1990–1996 and in columns (5)–(6) for the bust period 2007–2009. In columns (7)–(8) the sample is women aged 20–44 in the American Communities Survey for the bust period 2007–2009. The dependent variable in columns (3)–(8) is an indicator for having a child under one. In columns (1)–(2) ownership rates are matched by MSA, age category and race/ethnicity. In columns (3)–(8), ownership is the household's home ownership status, which is assigned as a 1 when the household owns a home and the respondent is the household head or spouse of the household head. House prices (10,000 s), unemployment rates, and male wages are matched by MSA and year. Columns (1)–(2) include group, MSA and year fixed effects. Columns (3)–(8) include fixed effects for race/ethnicity, age category and education. The instrumental variable is the Saiz (2010) supply elasticity measure interacted with the national version of the FHFA house price index. F Stat 1 refers to the first stage F statistic where the dependent variable is $HousePrice_{mt-1} * OwnRate_{mg}$, and F Stat 2 for $HousePrice_{mt-1}$. Data sources are: Census and Federal Housing Finance Agency (house prices), Current Population Survey (wages), and Bureau of Labor Statistics (unemployment rates). Standard errors adjusted for clustering at the MSA level are in parentheses. * $p<.1$, ** $p<.05$, *** $p<.01$.

An interesting empirical exercise is to consider the relative impact of unemployment rates versus housing prices. Using the same simulation procedure described above, we estimate the relative impacts of a one standard deviation increase in housing prices and decrease in unemployment rates. We find that at the mean rate of ownership (44%), a one standard deviation increase in housing prices leads to a 8.3% increase in births while a one standard deviation increase in unemployment rates leads to only a 2.1% decrease across all rates of ownership (note that this estimate is based on the point estimate in Table 4, column 5). Even among renters, the negative price effect an increase in housing prices is 21%, 10 times as large as the effect of unemployment rates. This highlights the importance of considering housing markets in any empirical analysis of how economic conditions affect fertility outcomes.[23]

5. Conclusion

This paper has investigated how current house prices affect current period birth rates. Our results suggest that house prices are a relevant factor in a couple's decision to have a baby at the present time. House prices lead to a negative price effect that conditionally reduces birth rates in the current period, and an offsetting positive home equity effect that leads to a net increase in births among homeowners. We use the estimated coefficients from our regression analyses to simulate the effect of a $10,000 increase in house prices on current year births. This exercise indicates that when home ownership rates reach 30%, the net effect becomes positive. At the mean U.S. home ownership in our sample period, the net effect of a $10,000 increase in prices is a 0.8% increase in births. Given underlying differences in home ownership rates and heterogeneity in the point estimates, the predicted net effect of house price changes varies across race/ethnic groups. We simulate that a $10,000 increase in MSA-level house prices leads to a 0.7% increase in current year births among Whites, a 0.2% increase in births among Blacks, and a 0.2% decrease in births among white Hispanics. Interestingly, these effects are substantially larger than the effects of changes in the unemployment rate. Moreover, using our estimates to make an out-of-sample prediction of the impact of the "Great Recession", we find that the fall in housing prices between 2006 and 2010 was associated with a 7.5% decline in births.

Our paper is written within the paradigm of the empirical literature on the cyclicality of fertility and as such, it is about the timing of fertility decisions. The evidence presented in our paper suggests that couples use some of their increased housing wealth to "fund" their childbearing goals. Our paper potentially demonstrates empirically that (imperfect) credit markets affect fertility timing. We have discussed our results in terms of the decision of whether or not to have a baby in the current period. We leave it to future research to investigate how house prices affect completed fertility or the demand for children more generally. In addition, it might also be true that when house prices increase or decrease, parents increase (or decrease)

[23] Schaller (2011) provides the most up to date and arguably compelling empirical analysis of the relationship because of local area unemployment rates and current period birth rates query. Her analysis finds that a one percentage-point increase in unemployment rates is associated with a 0.7 to 2.5% decrease in birth rates, depending on specification. Our specification finds that a one percentage-point increase in unemployment rates is associated with a 1.37% decrease in birth rates, which is well within her range of estimates. It is this estimated coefficient that we translate into a standard deviation measurement to compare to the effect of house prices. Therefore, our conclusion that house prices exert a larger effect on birth rates than do unemployment rates would apply even if we took an estimate of the cyclicality of birth rates from outside our own analysis.

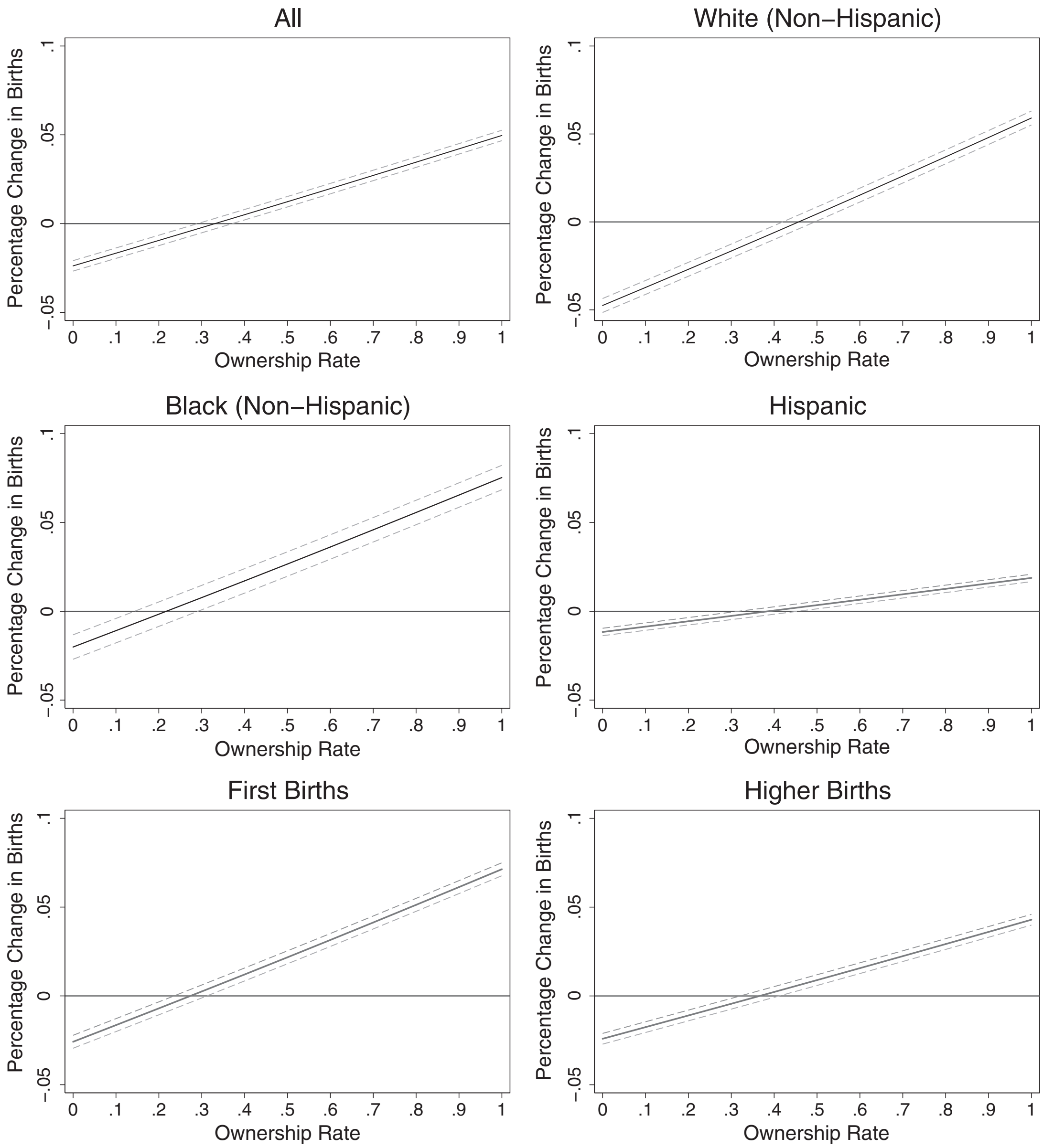

Fig. 3. Predicted percentage change in births for a $10,000 increase in MSA housing prices. Notes: These figures display the results of simulation exercises using estimates from the group specific IV regression specifications displayed in Table 7. We predict the percentage change in fertility rates from a $10,000 increase in mean housing prices for each ownership rate o displayed on the x axis: $(FertRate|HousePrice = h + 10k, OwnRate = o) - (FertRate|HousePrice = h, OwnRate = o)/(FertRate|HousePrice = h, OwnRate = o)$. For each group, we calculate the standard error of the prediction at the mean of the independent variables using 100 bootstrap replications and apply that standard error to calculate the confidence interval at each level of o. The solid line represents the predicted effect and the dashed line represents a 95% confidence interval, both of which were smoothed using a locally weighted linear regression.

quality investments in children, where quality of children is meant in the Beckerian sense. For example, perhaps some home-owning parents use their increased home equity to purchase, say, private education for their children. Once we allow for this possibility, it becomes clear that our empirical analysis is not designed to capture the full range of how real estate markets might affect childbearing and child rearing decisions.

Appendix A

A.1. Metropolitan areas

Metropolitan statistical areas are defined by the Office of Management and Budget. Their geographic definitions are based on core urban areas with a population of 50,000 or more and adjacent counties with a "high degree of social and economic integration (as measured by commuting to work) with the urban core" (Census Bureau documentation). Current metropolitan area definitions include both metropolitan areas (MSA) and divisions (MSADs), which are smaller units within this metropolitan area. Current definitions also include an alternative to the MSA/MSAD for metropolitan areas in the New England states, which are called New England City Town Areas (NECTAs). The boundaries of MSAs change over time as city populations change. The Office of Management and Budget releases revised definitions based on the decennial census and yearly census population estimates, and in addition to changing MSA compositions, sometimes changes the labels associated with each type of unit. A major change was done in 2003, at which point the coding system changed from a 4-digit coding system to a 5-digit coding system. Prior to 2003, instead of

MSADs Primary Metropolitan Statistical Areas (PMSAs) were used and instead of NECTAs, New England Metropolitan County Areas (NECMAs) were used.

The housing price index is available at the level of MSA/MSAD, based on the November 2008 definitions (released in December 2009), Since the index is based on repeat sales of the same home, the 2009 definitions apply throughout the data. For example, suppose a home sells once in 1980, 1990, and 2005. Suppose that in 1980 and 1990 it was not in an MSA, but in 2005 it was. Then, the home is considered part of the MSA and the housing price indices for 1980 and 1990 are revised to reflect the current boundaries. The rest of this Appendix A explains how we harmonize all other data sources to match this level of aggregation. Table 1 lists the level of geographic detail available for each of our control variables.

A.1.1. County level data

Whenever county level data is available, it is the preferred level of disaggregation because we can use it to construct MSA/MSADs which will exactly match the housing price index data. Data available at this level of disaggregation includes the Vital Statistics Natality Data (confidential files), Vital Statistics population data, Census median home value data, Bureau of Labor Statistics Unemployment data, Bureau of Economic analysis income per capita data, and the Rappaport and Sachs' (2003) coastal measure. To construct MSAs from the county level data, we use the 2009 metropolitan area definition files available from the Census Bureau at: http://www.census.gov/population/metro/files/lists/2009/List1.txt. These files map entire counties to a 2009 OMB MSA/MSAD definitions, thus, we can construct MSAs/MSADs that are exactly equivalent to those used in the housing price data.

It is worth noting a few technical points about linking counties to MSAs. First, Miami-Dade County, FL was renamed between the 1990 and 2000 census; so in all cases we have assigned the post-2000 FIPS code to this county.[24] Another issue concerns BLS Local Area Unemployment (LAU) Statistics, which are calculated at the county level, but use a coding system based on what are called “areas”. For the most part, the area codes are simply county FIPS codes. However, for counties which had large populations (50,000–100,000 and 100,000 plus) in 1970; a different coding system is applied.[25] We construct a crosswalk between the two using state FIPS codes and county names using vintage 2009 county FIPS codes.[26] Finally, in the BEA personal income data, BEA combines some counties/county equivalents in Virginia and assigns new county codes. We re-assign those counties which are contained within an MSA to one of the combined counties' FIPS code. In all cases these combinations were wholly contained within one MSA/MSAD.[27]

A.1.2. Vintage metropolitan area level data

For the case when counties are not available, but vintage metropolitan area definitions are available, we use those. By vintage metropolitan area definitions, we are referring to metropolitan areas based on historical definitions which may differ in composition from the 2009 definitions. Data that is available in this manner includes the 1990 and 2000 Census microdata (used to construct home ownership rates), the Current Population Survey data (used to construct wages and fraction college educated), and the Saiz (2010) elasticity measure. The vintage definitions used in these data include the 1983 MSA/PMSA, 1993 MSA/PMSA, 1999 MSA/NECMA, and 2003 MSA/NECTA codes, as described in Table 1.

To match the vintage definitions to the 2009 definitions, we begin by creating a crosswalk that links the counties that make up the different metropolitan areas over time. Unlike the current 2009 MSA/MSAD definitions (and vintage 2003 MSA/MSAD definitions) which directly map entire counties to MSAs, the earlier metropolitan area (and NECTA/NECMA) definitions allow for a single county to be in multiple metropolitan areas. For the case when a single county is in multiple MSAs/PMSAs/NECTAs/NECMAs, we use 1990 population counts of the minor civil divisions (a smaller unit within the metropolitan area) to assign the county to whichever MSAs/PMSAs/NECTAs/NECMAs the majority of the population resides.

From this county-MSA crosswalk, we construct vintage MSA-to-2009 MSA/MSAD crosswalks. In most cases, there is a one on one match between the vintage MSA definitions and the 2009 definitions. In some cases, however, it's possible for a vintage metropolitan area to have split into two or combined to form a single metropolitan area by 2009. For metropolitan areas that have combined to form one metropolitan area by 2009, we use 1990 population weights to create a population weighted average of the data. For metropolitan areas that have split, we apply the single data point to all the split-off areas.

A.1.3. Attaching aggregate measures to the individual level data

In the individual level data, we are given the vintage metropolitan area codes. In this case, we need to construct the housing price, wage, and unemployment data according to those definitions. In the individual CPS we are provided with 1983, 1993 and 2003 MSA/2003 NECTA codes and in the AHS we are provided with 1980 MSA codes. For the ACS, only PUMAs (Public Use Microdata Areas) are provided, however, IPUMS has created a crosswalk procedure and attached 1993 MSA codes, which we will use Ruggles et al. (2010). Recall that the unemployment data is at the county level. In this case, we use county-to-vintage MSA cross walk described in the section above. For the wage data, linking to the CPS is trivial since it was constructed in the CPS and therefore uses the same MSA definitions. For linking the wage data to the ACS and for linking the housing data to the CPS, ACS and AHS, we again use the county-to-vintage MSA crosswalk described above. In this case, if multiple 1980/1983/1993/2003 MSA/2003 NECTA combines to form a single MSA in 2009, we assign the housing price data to each vintage MSAs. For the case when a single 1980/1983/1993/2003 MSA/2003 NECTA splits to form multiple MSAs in 2009, we use 1990 population weights to assign a weighted average of home prices to the vintage metropolitan areas codes. Finally, since CPS uses different MSA codes over time which are not consistent, we use the linked 2009 MSA definition for the fixed effects. In the case where the vintage MSA splits into multiple 2009 MSADs, we use assign the code of the MSAD with the largest population share.

A.2. Construction of house prices

We use the same procedure used by Glaeser et al. (2008) to construct house prices. First, we construct a 2000 median home value from county-level census data, using the crosswalk procedure outlined above to create a population-weighted median home value. We inflate this value to 2006 dollars using the CPI-U “All Items-Less Shelter Series.” We then take this value and scale it by the percent change in the housing price index from 2000 to the year of interest, which is calculated: $(hpi_t - hpi_{2000})/hpi_{2000}$. The housing price index is also inflated to 2006 dollars using the CPI-U “All Items-Less Shelter Series” prior to scaling. This gives us a value that proxies for the price growth of a median value home in each MSA over time.

[24] See, for example, http://www.census.gov/popest/archives/files/90s-fips.txt.
[25] See http://www.bls.gov/lau/laucodes.htm.
[26] http://www.census.gov/popest/geographic/codes02.html.
[27] See http://www.bea.gov/regional/docs/msalist.cfm.

Appendix Table 1
Characteristics of metropolitan areas in the sample.

Metropolitan area name (2009 MSAD)	Percent change prices 97-06	Home price 2006	Elasticity of supply	Unemployment rate	Median wage 2006
Salinas, CA	171.4%	$588,736	1.10	6.92	$15.83
Santa Barbara-Santa Maria-Goleta, CA	165.5%	$628,696	0.89	4.04	$12.02
Riverside-San Bernardino-Ontario, CA	162.7%	$338,547	0.94	4.92	$16.83
Los Angeles-Long Beach-Glendale, CA	162.6%	$515,061	0.63	4.78	$16.83
Vallejo-Fairfield, CA	151.9%	$398,870	1.14	4.87	$18.27
San Diego-Carlsbad-San Marcos, CA	151.4%	$474,242	0.67	3.96	$19.23
Oxnard-Thousand Oaks-Ventura, CA	148.7%	$556,284	0.75	4.30	$19.23
Fort Lauderdale-Pompano Beach-Deerfield Beach, FL	146.5%	$267,370	0.65	3.07	$15.11
Stockton, CA	146.4%	$331,468	2.07	7.42	$21.31
Modesto, CA	145.8%	$314,814	2.17	7.96	$18.69
Oakland-Fremont-Hayward, CA	144.1%	$564,108	0.70	4.37	$24.04
Miami-Miami Beach-Kendall, FL	142.6%	$295,201	0.60	4.08	$15.11
West Palm Beach-Boca Raton-Boynton Beach, FL	140.7%	$290,506	0.83	3.64	$15.11
Santa Rosa-Petaluma, CA	131.4%	$510,412	1.00	3.99	$28.00
Cape Coral-Fort Myers, FL	131.3%	$241,947	1.28	2.88	$17.79
Fresno, CA	126.2%	$264,992	1.84	8.01	$16.17
North Port-Bradenton-Sarasota, FL	125.7%	$247,504	0.92	3.06	$19.40
Port St. Lucie, FL	125.4%	$238,298	1.19	3.89	$17.55
Bakersfield-Delano, CA	123.9%	$230,694	1.64	7.54	$18.33
San Francisco-San Mateo-Redwood City, CA	122.9%	$781,891	0.66	3.89	$24.04
Deltona-Daytona Beach-Ormond Beach, FL	117.7%	$194,512	1.07	3.24	$16.83
Washington-Arlington-Alexandria, DC-VA-MD-WV	117.2%	$401,637	1.61	3.13	$23.08
Palm Bay-Melbourne-Titusville, FL	117.0%	$210,691	1.04	3.23	$17.55
San Jose-Sunnyvale-Santa Clara, CA	116.9%	$698,468	0.76	4.55	$24.04
Tampa-St. Petersburg-Clearwater, FL	112.7%	$186,334	1.00	3.43	$19.23
Orlando-Kissimmee-Sanford, FL	107.0%	$221,037	1.12	3.12	$15.42
Bethesda-Rockville-Frederick, MD	106.4%	$445,442	1.61	2.87	$23.08
Phoenix-Mesa-Glendale, AZ	106.4%	$252,298	1.61	3.62	$16.35
Atlantic City-Hammonton, NJ	105.1%	$258,536	1.12	5.68	$23.32
New York-White Plains-Wayne, NY-NJ	104.7%	$472,658	0.80	4.79	$21.63
Providence-New Bedford-Fall River, RI-MA	98.4%	$272,732	1.34	5.37	$13.10
Visalia-Porterville, CA	94.6%	$224,105	1.97	8.50	$16.83
Boston-Quincy, MA	94.2%	$349,966	0.86	4.66	$25.56
Poughkeepsie-Newburgh-Middletown, NY	94.0%	$290,782	1.79	4.12	$23.61
Las Vegas-Paradise, NV	92.6%	$288,093	1.39	4.28	$17.55
Jacksonville, FL	92.6%	$181,653	1.06	3.26	$19.62
Baltimore-Towson, MD	89.8%	$258,936	1.23	4.00	$23.45
Ocala, FL	89.2%	$147,002	1.73	3.39	$12.02
Newark-Union, NJ-PA	88.6%	$381,183	1.17	4.65	$21.63
Charleston-North Charleston-Summerville, SC	85.6%	$176,563	1.20	5.10	$19.23
Virginia Beach-Norfolk-Newport News, VA-NC	85.5%	$221,630	0.82	3.32	$16.68
Reno-Sparks, NV	84.7%	$316,406	1.39	4.12	$16.83
Lakeland-Winter Haven, FL	82.9%	$141,875	1.56	3.60	$15.68
Peabody, MA	81.8%	$335,451	0.86	5.09	$25.56
Bridgeport-Stamford-Norwalk, CT	81.0%	$468,745	0.98	3.90	$29.91
Trenton-Ewing, NJ	80.7%	$276,183	1.88	4.24	$23.02
Worcester, MA	80.6%	$245,583	0.86	5.10	$20.00
Seattle-Bellevue-Everett, WA	77.4%	$369,177	0.88	4.28	$23.94
Tucson, AZ	76.5%	$198,430	1.42	4.01	$14.42
Cambridge-Newton-Framingham, MA	76.0%	$380,977	0.86	3.94	$25.56
Gainesville, FL	75.6%	$169,875	2.48	2.66	$16.67
New Haven-Milford, CT	71.3%	$261,064	0.98	4.85	$23.62
Tacoma, WA	71.0%	$261,712	1.21	5.05	$23.94
Camden, NJ	70.5%	$232,665	1.65	4.68	$21.45
Norwich-New London, CT	68.2%	$252,431	1.46	4.17	$17.31
Philadelphia, PA	67.6%	$211,020	1.65	4.50	$21.45
Minneapolis-St. Paul-Bloomington, MN-WI	66.5%	$221,112	1.45	3.83	$22.12
Wilmington, DE-MD-NJ	63.9%	$232,676	1.95	3.92	$21.45
Pensacola-Ferry Pass-Brent, FL	61.2%	$154,225	1.48	3.07	$15.38
Vineland-Millville-Bridgeton, NJ	61.0%	$166,452	1.85	6.92	$19.23
Springfield, MA	59.4%	$209,209	1.52	5.20	$20.60
Richmond, VA	58.0%	$185,521	2.60	3.19	$19.23
Olympia, WA	57.3%	$248,438	1.75	4.56	$21.06
Hartford-West Hartford-East Hartford, CT	54.2%	$238,239	1.50	4.54	$24.04
Portland-Vancouver-Hillsboro, OR-WA	52.9%	$283,856	1.07	5.02	$20.41
Allentown-Bethlehem-Easton, PA-NJ	52.8%	$207,168	1.77	4.55	$19.23
Albany-Schenectady-Troy, NY	52.8%	$187,483	1.70	3.96	$18.63
Asheville, NC	51.8%	$159,848	1.55	3.75	$17.79
Chicago-Joliet-Naperville, IL	49.7%	$251,216	0.81	4.46	$20.66
Denver-Aurora-Broomfield, CO	40.1%	$216,292	1.53	4.46	$21.37
Milwaukee-Waukesha-West Allis, WI	38.2%	$180,640	1.03	4.85	$20.14
Spokane, WA	37.4%	$182,067	1.64	4.94	$15.87
York-Hanover, PA	36.8%	$172,685	1.99	3.97	$21.15
St. Louis, MO-IL	35.9%	$136,093	2.36	5.09	$19.23
Lake County-Kenosha County, IL-WI	35.4%	$258,459	1.00	4.65	$20.66
Racine, WI	35.1%	$161,554	1.77	5.61	$20.19

Appendix Table 1 (*continued*)

Metropolitan area name (2009 MSAD)	Percent change prices 97-06	Home price 2006	Elasticity of supply	Unemployment rate	Median wage 2006
Madison, WI	34.0%	$201,667	2.25	3.41	$19.23
Reading, PA	33.9%	$166,269	2.03	4.34	$20.61
Fayetteville-Springdale-Rogers, AR-MO	32.7%	$132,146	2.06	3.55	$16.33
Austin-Round Rock-San Marcos, TX	32.7%	$157,418	3.00	4.15	$19.79
Lancaster, PA	31.8%	$180,168	2.24	3.47	$19.23
Binghamton, NY	31.7%	$107,227	2.26	4.65	$20.43
Houston-Sugar Land-Baytown, TX	30.9%	$113,243	2.23	5.00	$15.87
Utica-Rome, NY	30.1%	$106,208	2.79	4.57	$9.68
Colorado Springs, CO	30.0%	$198,587	1.67	4.69	$17.09
Atlanta-Sandy Springs-Marietta, GA	29.5%	$179,457	2.55	4.63	$18.13
Albuquerque, NM	28.2%	$183,194	2.11	3.92	$16.33
Mobile, AL	27.9%	$107,656	2.04	3.60	$22.24
Niles-Benton Harbor, MI	27.8%	$127,136	2.06	6.89	$23.56
Kansas City, MO-KS	27.6%	$132,024	3.19	5.03	$20.31
Salt Lake City, UT	27.4%	$225,663	0.75	2.96	$17.09
Lafayette, LA	27.1%	$120,141	4.84	2.89	$18.03
Baton Rouge, LA	27.0%	$122,634	1.74	3.93	$20.03
Ann Arbor, MI	26.6%	$212,388	2.29	4.56	$35.26
Chattanooga, TN-GA	26.4%	$119,291	2.11	4.40	$16.35
El Paso, TX	25.9%	$99,742	2.35	6.71	$13.46
Knoxville, TN	25.8%	$131,259	1.42	4.15	$19.51
Lexington-Fayette, KY	25.8%	$140,213	2.63	4.63	$19.90
Birmingham-Hoover, AL	24.7%	$121,813	2.14	3.22	$19.23
San Antonio-New Braunfels, TX	24.6%	$102,182	2.98	4.61	$14.05
Syracuse,NY	24.5%	$116,321	2.21	4.69	$19.44
Corpus Christi, TX	24.1%	$94,566	1.65	4.95	$14.79
Harrisburg-Carlisle, PA	23.6%	$150,613	1.63	3.67	$18.38
Nashville-Davidson–Murfreesboro–Franklin, TN	23.2%	$161,298	2.24	4.23	$18.49
Augusta-Richmond County, GA-SC	22.9%	$109,720	3.57	5.85	$20.09
Columbia, SC	22.8%	$119,656	2.64	5.53	$14.42
Lansing-East Lansing, MI	22.4%	$133,123	2.58	5.79	$19.71
Scranton–Wilkes-Barre, PA	22.2%	$123,566	1.62	5.15	$15.82
Detroit-Livonia-Dearborn, MI	21.4%	$113,685	1.24	8.39	$24.04
Des Moines-West Des Moines, IA	20.0%	$129,175	3.66	3.38	$19.23
Davenport-Moline-Rock Island, IA-IL	19.4%	$105,763	4.11	4.28	$16.44
Durham-Chapel Hill, NC	19.4%	$171,002	2.11	3.91	$16.49
Dallas-Plano-Irving, TX	19.1%	$124,055	2.18	4.82	$18.73
Little Rock-North Little Rock-Conway, AR	19.0%	$108,394	2.79	4.68	$14.74
Louisville/Jefferson County, KY-IN	18.6%	$127,502	2.34	5.67	$16.02
Pittsburgh, PA	18.3%	$110,200	1.20	4.68	$19.15
Warren-Troy-Farmington Hills, MI	18.2%	$186,579	1.30	6.42	$24.04
Beaumont-Port Arthur, TX	18.1%	$76,356	2.49	5.89	$14.62
Grand Rapids-Wyoming, MI	16.5%	$131,715	2.39	5.80	$18.18
Charlotte-Gastonia-Rock Hill, NC-SC	16.3%	$151,494	3.09	4.76	$17.20
Omaha-Council Bluffs, NE-IA	16.2%	$124,335	3.47	3.45	$19.23
Kalamazoo-Portage, MI	15.7%	$127,799	2.48	5.45	$23.08
Fort Worth-Arlington, TX	15.7%	$108,505	2.80	4.72	$18.73
Hickory-Lenoir-Morganton, NC	15.6%	$105,663	2.41	5.85	$16.15
Cincinnati-Middletown, OH-KY-IN	15.5%	$138,971	2.51	5.14	$21.18
Lubbock, TX	15.0%	$83,529	4.33	3.98	$14.66
Peoria, IL	14.4%	$110,385	3.23	4.17	$16.08
Columbus, OH	13.8%	$146,041	2.71	4.65	$19.23
Rockford, IL	13.8%	$123,510	3.68	5.63	$16.83
Toledo, OH	13.8%	$114,567	2.21	5.99	$24.55
Raleigh-Cary, NC	13.7%	$178,667	2.11	3.70	$21.45
Flint, MI	13.6%	$108,328	2.75	8.04	$19.23
Greenville-Mauldin-Easley, SC	13.5%	$117,495	2.71	5.64	$16.83
Gary, IN	13.2%	$131,597	1.74	5.35	$20.66
Springfield, MO	13.1%	$114,971	3.60	3.87	$16.24
South Bend-Mishawaka, IN-MI	12.8%	$107,023	4.36	5.14	$17.63
Winston-Salem, NC	12.4%	$128,671	3.10	4.27	$13.22
Memphis, TN-MS-AR	11.8%	$108,828	1.76	5.68	$19.71
Wichita, KS	11.6%	$96,868	5.45	4.57	$16.35
Montgomery, AL	11.1%	$108,348	3.58	3.45	$14.90
Ogden-Clearfield, UT	10.9%	$179,163	0.75	3.14	$22.84
Saginaw-Saginaw Township North, MI	10.9%	$100,272	2.23	7.35	$14.42
Buffalo-Niagara Falls, NY	9.9%	$115,101	1.83	5.11	$19.23
Greensboro-High Point, NC	9.8%	$122,924	3.10	4.79	$16.83
Akron, OH	9.0%	$132,057	2.59	5.16	$18.03
Cleveland-Elyria-Mentor, OH	8.7%	$140,618	1.02	5.52	$19.47
Canton-Massillon, OH	8.6%	$117,865	3.03	5.70	$15.87
Fayetteville, NC	7.8%	$107,010	2.71	5.38	$14.42
Youngstown-Warren-Boardman, OH-PA	7.5%	$97,192	2.59	6.07	$18.96
Rochester, NY	7.3%	$117,940	1.40	4.56	$19.23
Indianapolis-Carmel, IN	7.0%	$130,473	4.00	4.36	$19.23
Spartanburg, SC	6.8%	$100,563	2.71	6.59	$16.75
Dayton, OH	3.9%	$119,761	3.71	5.65	$17.07
Fort Wayne, IN	3.3%	$101,869	5.36	4.89	$19.23

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