DO FAMILY WEALTH SHOCKS AFFECT FERTILITY CHOICES? EVIDENCE FROM THE HOUSING MARKET

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Abstract—This paper uses wealth changes driven by housing market variation to estimate the effect of family resources on fertility decisions. Using data from the Panel Study of Income Dynamics, we show that a $100,000 increase in housing wealth among home owners causes a 16% to 18% increase in the probability of having a child. There is no evidence of an effect of MSA-level housing price growth on the fertility of renters, however. We also present evidence that housing wealth growth increases total fertility and that the responsiveness of fertility to housing wealth has increased over time, commensurate with the recent housing boom.

I. Introduction

A long literature in economics dating back to Malthus (1798) has focused on the fertility decisions of households and how these decisions are influenced by financial incentives. Becker (1960) introduced children into economic models as a durable good in the utility function of the parents. Because there are few substitutes for children, they generally are assumed to be normal goods, implying that fertility should respond positively to an increase in household income or wealth.

The empirical evidence has been largely inconsistent with this assumption. Across countries, there is a strong negative correlation between GDP and fertility. Within countries, there is cross-sectional evidence of a negative correlation between income and fertility across households (see Jones, Schoonbroodt, & Tertilt, 2010). For example, in the United States, Jones and Tertilt (2008) estimate an income elasticity for fertility of about −0.38 using data from the U.S. Census over the past century and a half. Time series data yield similar findings: household income has increased while fertility rates have declined. A large number of papers also link the higher incomes that came with the industrial revolution to the demographic transition of industrial countries (examples include Clark, 2005; Galor, 2005; Bar & Leukhina, 2010).

Becker (1960) assumed that children were normal goods, but then had to reconcile this assumption with the observed negative correlation between income and fertility. He added child quality to his model, which created a quantity-quality trade-off, to generate the negative relationship between income and fertility.[1] While there is a strong negative correlation between family size and various measures of child quality, including education and future earnings, this correlation does not mean that the quantity-quality trade-off is causal. Angrist, Lavy, and Schlosser (2010) find no evidence of a quantity-quality trade-off using exogenous variation in family size due to twin births and sibling-sex composition. This finding highlights the puzzle of what explains the strong negative correlation between income and fertility if there is no evidence of a trade-off between child quantity and quality.

One potential solution to this puzzle is that the negative correlation between fertility and income in the cross-section does not reflect a causal relationship. The opportunity cost of female time could be a contributing factor to the bias associated with cross-sectional estimates. Butz and Ward (1979), Schultz (1985), and Heckman and Walker (1990) all find evidence that fertility is decreasing in female wage rates and argue that the negative correlation between income and fertility is a substitution effect due to higher wages, not an income effect. However, even controlling for female wages, the negative correlation between income and fertility generally remains.

Differences in cost of living across areas also could generate a negative cross-sectional relationship between fertility and income. Higher cost of living implies a higher cost of goods that are complementary to raising a child, such as housing, food, and education. As shown in panel A of figure 1, the fertility rate is lower in states where housing is more expensive. Panel B depicts the negative correlation between income and fertility. Panel C demonstrates the difficulty in interpreting the evidence in panels A and B as causal. This panel shows a strong positive relationship between housing prices and income, which means that the places with the highest incomes also have the highest cost of living. High-income households may have low cost-of-living-adjusted real incomes that could lower their fertility. Additionally, households with lower underlying fertility may select into areas with higher housing prices and thus a higher cost of raising a child. Such selection implies that the cross-sectional estimates of the effect of income and housing price on fertility potentially contain large negative biases.[2]

Several recent studies have sought to overcome these problems by using arguably exogenous income shocks to identify the income effect on fertility. Lindo (2010) and Amialchuk (2006) both show that fertility is negatively affected by shocks

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[1] Jones, Schoonbroodt, and Tertilt (2010) show that adding a quality choice by itself does not generate a negative income-fertility relationship without also assuming a high elasticity of substitution between children and consumption.

[2] Black et al. (forthcoming) make a similar point.

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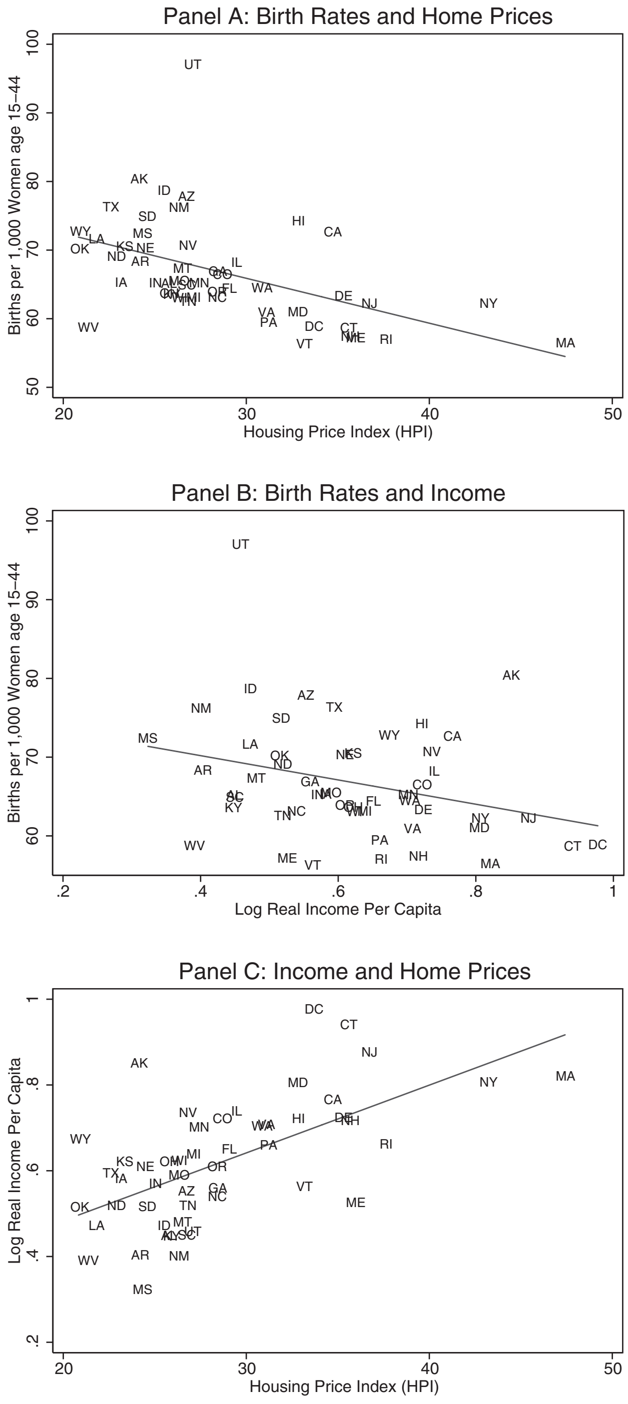
FIGURE 1.—CROSS-SECTIONAL RELATIONSHIPS BETWEEN STATE BIRTH RATES, REAL HOUSING PRICES, AND LOG PER CAPITA INCOME, 1976–2008



Birth rates are calculated from the number of births per thousand women ages 15–44 in each state, reported quarterly by the CDC National Vital Statistics Reports from 1976 to 2008, divided by the yearly population of women ages 15–44 in each state as reported by the U.S. Census Bureau. State-level home prices come from the state-level Federal Housing Finance Agency Housing Price Index (HPI). Per capita income is collected from the Bureau of Labor Statistics. All financial variables are inflated to 2008 dollars using the CPI-U.

to family income brought about by male job loss. However, these authors face difficulties in disentangling the effect of job loss per se on fertility from the effect on family income. Black et al. (forthcoming) find that the 1970s West Virginia coal boom created a large positive shock to male income and subsequently increased fertility. The restricted geographic region used in this study introduces concerns about the generalizability of the results to the rest of the United States, though. While all three papers provide evidence that fertility and family resources are positively linked, the primary identification issue for each is that a change in wage affects the opportunity cost of time. Focusing only on male wage shocks does not fully alleviate the concern, as male wages also may alter the allocation of household work between husband and wife, which likely contributes to the decision to have a child.

In this paper, we use family wealth variation supplied by the housing market to identify how household resources affect fertility decisions. Our analysis makes several contributions to the existing literature. First and foremost, our use of home price variation allows us to identify the effect of household resources on fertility using variation that has no first-order effect on the opportunity cost of time or on time allocation between market and home production and that overcomes the biases associated with using cross-sectional data. Our analysis also uses housing variation from a more nationally representative sample than these other papers, which makes the estimates more generalizable than some of the previous work that focuses on a particular region or the availability of plant closings. Furthermore, this paper is the first to examine how household fertility responds to the wealth of the household rather than simply its income. Excluding wealth may be particularly problematic because it can cause one to mischaracterize the financial resources of the household. If fertility decisions are a function of total resources, using income as a proxy may yield an incomplete picture of how resources affect fertility.

We analyze housing wealth as our measure of household wealth for several reasons. First, about 54% of women of childbearing age in the Panel Study of Income Dynamics (PSID), the data we use in this analysis, own a home. Second, for these women, and for the United States as a whole, housing wealth represents the vast majority of total household wealth.[3] Third, we argue that housing market changes can be used to generate exogenous household resource variation, which allows us to overcome the inherent endogeneity between wealth accumulation and childbearing decisions. The main identifying variation in this analysis comes from the housing boom that began in the late 1990s and was characterized by large increases in home prices that occurred differentially across cities and an increased liquidity of home equity. Home owners who lived in high-growth areas experienced a large increase in their liquid wealth relative to home owners in lower-growth areas and relative to renters throughout the United States. In examining the effect of housing wealth changes on fertility decisions, our paper thus contributes to the growing literature on housing wealth

[3] For example, in the 2004 survey of consumer finances, home equity accounted for 85% of household wealth for the median home owner between 25 and 55 years old. Median nonhousing wealth among this group was $11,661, and the bottom quartile had less than $2,500 in nonhousing wealth.

and household behavior as well. There is a debate among economists as to whether housing wealth has any effect on savings, consumption, education, and labor supply decisions because housing wealth has traditionally been hard to realize without selling the home. This analysis provides evidence that changes in housing wealth influence fertility, which suggests that it also may affect other household decisions.

Using restricted-access PSID data from 1985 to 2007 that contain geographic identifiers, we use short-run home price variation over time within cities to examine whether families in higher-growth areas made different fertility decisions from families in lower-growth areas, controlling for detailed demographic characteristics as well as for city fixed effects. Lovenheim (2011) uses a similar method for identifying the causal effect of housing wealth on college attendance, and it is based on the assumption that the geographic variation in the strength and timing of the housing boom was conditionally exogenous to individual household behavior. Our main finding is that short-run increases in one's home value is associated with a positive and significant increase in the likelihood of having a child. The marginal effects appear small: a \$100,000 increase in the value of a home over the prior four years is associated with 0.85 of a percentage point increase in the likelihood of having a child. However, the baseline probability of having a child is 5.0% for women age 25 to 44 who own a home, so an increase of 0.85 represents a 16.4% increase in the probability of having a child. This implies a housing wealth elasticity of about 0.13. We also show that only home owners experience the positive wealth shock from the housing boom, while both home owners and renters experience the substitution effect from the increase in housing costs. Among renters, housing price increases have little effect and may even reduce the likelihood of giving birth, which suggests our estimates are not being driven by unobserved economic shocks at the state or local level.[4]

This analysis also contributes to existing work on family resources and fertility by identifying the effect of housing wealth on total fertility using the age distribution of respondents. We document sizable effects of home price changes on the likelihood of birth among women of all ages between 25 and 45, suggesting that wealth shocks increase the total amount of fertility rather than just its timing. Our estimates also point to reductions in the birth probability for 15 to 19-year-old women when the housing wealth of their parents increases. This evidence points to reductions in teen births when household resources increase. Our empirical analysis concludes by examining heterogeneity by number of children already born, family income, and time period.

[4] Using state-level fertility and housing price data, we find a positive and significant relationship between housing wealth and fertility. These results are consistent with a positive wealth shock causing an increase in fertility and are large despite the fact that only about half of women of child-bearing age own homes. However, we cannot separate home owners from renters using state-level aggregate data. In addition, the identification is susceptible to the claim of an unobserved correlated shock at the state level that is driving both fertility decisions and home prices (such as unobserved economic shocks).

II. Children and Housing

Children and housing have intrinsic characteristics that make them different from other goods. Focusing first on children, we follow Becker (1960) and model families as choosing how many children to have in a utility-maximizing framework. Families raise children using a combination of parent time and market goods. The cost of raising a child differs over families and depends on the value of parent time, the cost of relevant market goods, and the function form of the child production function. One of the market goods that belongs in the child production function is housing. Thus, changes in home prices may affect the price of raising a child.

More formally, let N denote the number of children a family chooses to have and P_N be the cost of raising an additional child. Similarly, let the amount of housing be given by H and the price of housing be given by P_H. The optimal quantities of N and H are determined by the usual marginal utility condition:

$$\frac{U_N}{U_H} = \frac{P_N}{P_H}. \tag{1}$$

Just as an increase in the value of parent time would cause an increase in P_N, an increase in P_H leads to an increase in P_N to the extent that housing is an input in the child production function. Thus, increases in home prices induce a potential substitution effect, whereby families substitute away from complements to housing (such as children) to other goods. At the same time, when home prices rise, this price increase is capitalized into housing wealth for a home owner. If children are normal goods, this wealth increase should cause families to increase fertility. The implication of the substitution and wealth effects for our estimates that use P_H as a source of variation in household wealth is that they likely are biased toward 0 because any positive effect has embedded in it a negative substitution effect. The size of this bias clearly depends on how much P_N increases due to an increase in home prices, which neither we nor previous analyses have been able to estimate.[5] Given that our estimates are positive and large relative to baseline fertility rates, we do not believe this bias is a serious concern, and it suggests that the true income effect may be even larger than our estimates indicate. Furthermore, the similarity of our estimates to those that use male labor force separation as an instrument for income (Lindo, 2010;

[5] There is evidence that fertility responds positively to exogenous changes to the financial incentives of having a child (a reduction in P_N). Baughman and Dickert-Conlin (2003) find some evidence of a positive fertility response to earned income tax credit (EITC) expansions in the 1990s, primarily for married nonwhite women. Crump, Goda, and Mumford (2011) show that the long-run effect of child tax benefits in the United States on fertility is small but positive, primarily operating through the timing of births. Internationally, Milligan (2005) finds a large fertility response to a temporary child subsidy program in Quebec; however, Parent and Wang (2007) show that women in Quebec may have had children earlier in order to claim the subsidy with no change in their completed fertility. Cohen, Dehejia, and Romanov (forthcoming) estimate a strong positive effect of financial incentives on fertility among low-income populations in Israel.

Amilachuk, 2006) supports both types of analyses because the bias stemming from the lower opportunity cost of male time when men lose their job is positive.

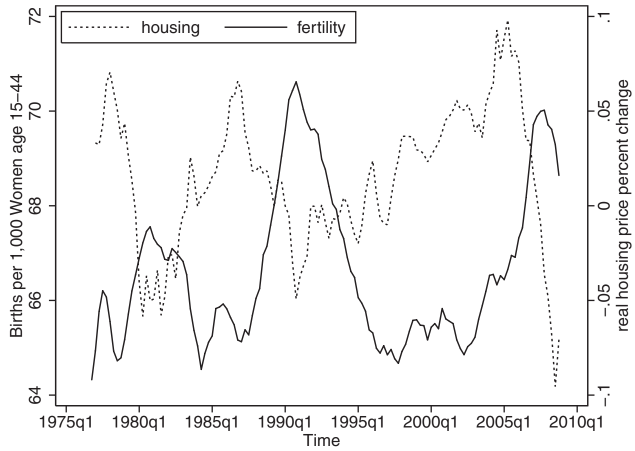
A critical factor in interpreting our estimates is that housing wealth must reflect real wealth. If not, then home price increases will cause only a substitution away from household-based goods such as children. For a home owner, if all home prices increase similarly in an area, then the home owner is potentially no wealthier because it would be very hard to realize the wealth increase since the price of all housing consumption has risen similarly. While it would be possible for a home owner to sell her house and move into a less expensive home or to an area that did not experience the price increase, such moves are rare, particularly among families on the margin of having more children. Alternatively, the home owner can take out a home equity loan or line of credit, thereby enabling her to realize the increased wealth in the home without selling it. As Greenspan and Kennedy (2005) have documented, equity extraction became much more prevalent during the late 1990s and 2000s, which likely increased household responsiveness to any home price increases. We will present estimates by decade in order to determine the extent to which the increased liquidity of housing wealth affected the impact of home price changes on fertility. Because the increase in home equity liquidity should not alter the substitution effect generated by home price increases, this comparison also will lend insight into the role of liquid wealth per se on fertility decisions.

III. Data

We start by examining national fertility and housing time-series data in order to describe the aggregate correlation between home price variation and fertility. We construct a quarterly birthrate measure, which is the number of births per thousand women ages 15 to 44 from the CDC National Vital Statistics Reports, from 1976 to 2008.[6] Our housing price measure is the state-level Federal Housing Finance Agency Housing Price Index (HPI), averaged using the number of women ages 15 to 44 as the weights.[7] This index is constructed from all repeat-sale, single-family homes whose mortgages have been securitized by Fannie Mae or Freddie Mac in each year. In order to make the index comparable across several years, we scale it by the CPI-U to put it into constant 2008 dollars.

The correlation between housing price changes and birth rates over time is shown in figure 2. This figure plots the births per 1,000 women ages 15 to 44 in the country as the solid line; the dashed line is the percentage change in real housing prices. Although on different scales, these data are consistent with a negative cross-sectional relationship between housing price variation and fertility. In times of high housing price growth, fertility appears to fall, and vice versa. Thus, at least in the aggregate, this figure is suggestive that housing prices negatively affect fertility. However, figure 2 also shows evidence of a delayed positive fertility response to an increase in real home prices: fertility follows the housing price change with a two- to three-year lag. Though not strong evidence, it is suggestive of the need to consider short-run home price changes in our empirical specifications that use microdata.

Figure 2.—National Birth Rate and Real Housing Price Percentage Change, 1976–2008



Birth rates are calculated from the number of births per thousand women ages 15–44, reported quarterly by the CDC National Vital Statistics Reports from 1976 to 2008, divided by the yearly population of women ages 15–44 as reported by the U.S. Census Bureau. State-level home prices come from the quarterly state-level Federal Housing Finance Agency Housing Price Index (HPI), averaged across states using the population of women ages 15–44 in each year. The home price index is inflated to 2008 dollars using the CPI-U and is expressed as a quarterly growth rate.

The microdata we use for the analysis are the 1985–2007 restricted-use data from the PSID, a longitudinal data set that began with a representative set of households in 1968. Since that time, it has followed these respondents and their descendants continually and has added refresher samples that include recent immigrants in the 1990s. The main advantages of the PSID over other available survey data is that it is a long panel that allows us to track changes in the family's home price prior to a child's birth. The data also contain a rich set of individual and family background information that is instrumental in controlling for selection of families with different fertility patterns into cities with different housing growth rates. We use the restricted-use geocode files that identify the metropolitan statistical area (MSA), or city, in which each woman lives. These geographic identifiers allow us to control for cross-city selection correlated with unobserved fertility preferences of households.

Our sample is women ages 25 to 44 who are or are descended from an original PSID survey respondent.[8] We

[6] These data are available at http://www.cdc.gov/nchs/nvss.htm.

[7] These data are available at http://www.fhfa.gov. This index, previously known as the Office of Federal Housing Enterprise Oversight (OFHEO) index, is widely used in the housing literature.

[8] Many women appear in the sample because they marry or cohabitate with an original PSID member or descendant. If the relationship ends, the woman no longer is in the sample. Thus, we focus on the sample of women we can follow continuously over time to avoid sample selection biases driven by divorce and other breakups.

TABLE 1.—PSID SUMMARY STATISTICS FOR HOME OWNERS AND RENTERS

Variable	Mean	S.D.	Minimum	Maximum
A. Home Owners				
Birth	0.050	0.218	0	1
Home value ($100,000)	1.5853	1.4415	.0104	33.2274
2-year home value change ($100,000)	0.3496	0.7313	−2.9668	4.9609
4-year home value change ($100,000)	0.6072	0.9881	−2.8814	14.0415
Married	0.810	0.392	0	1
Real family income ($100,000)	0.8620	0.6910	−0.8142	24.0098
Children	1.539	1.194	0	9
Age 25–29	0.191	0.393	0	1
Age 30–34	0.259	0.438	0	1
Age 35–39	0.288	0.453	0	1
Age 40–44	0.262	0.440	0	1
High school drop out	0.115	0.319	0	1
High school diploma	0.388	0.487	0	1
Some college	0.244	0.429	0	1
College graduate	0.216	0.412	0	1
Education missing	0.037	0.189	0	1
B. Renters				
Birth	0.054	0.226	0	1
Market average home price ($100,000)	1.6367	0.8285	0.4212	6.9224
2-year market home value change ($100,000)	0.0664	0.4068	−2.8971	3.8146
4-year market home value change ($100,000)	0.1150	0.5927	−2.9074	7.5964
Married	0.457	0.498	0	1
Real family income ($100,000)	0.4051	0.3495	−.8652	15.5748
Children	1.560	1.386	0	9
Age 25–29	0.333	0.471	0	1
Age 30–34	0.291	0.454	0	1
Age 35–39	0.219	0.414	0	1
Age 40–44	0.157	0.364	0	1
High school drop out	0.241	0.428	0	1
High school diploma	0.382	0.486	0	1
Some college	0.223	0.416	0	1
College graduate	0.101	0.301	0	1
Education missing	0.052	0.222	0	1

For home owners, the number of observations is 32, 218; for renters, 27,252. In panel A, all home values are self-reported and apply to the home owner. In panel B, housing values are averages among home owners in the sample at the MSA-by-year level. We calculate means of these averages over all renters, which is what is shown in panel B. Negative income values indicate net losses for the family in that year. All monetary means are in real 2008 dollars and were inflated using the CPI-U.

Source: 1985–2007 Panel Study of Income Dynamics, women age 25–44; those who moved in the prior four years are excluded.

focus on this age group because the birth rate among women over 45 is extremely low and because women under 25 who live in a home that someone in their household owns typically live with their parents. Using the PSID natality files, which contain a detailed record of all births to sample participants beginning in 1985, we construct the variable *birth*, a dummy variable equal to 1 if a woman gives birth within a year prior to the survey date. We use the reported market value of the home as our home price measure. This value is reported by the respondent in each survey and thus is consistently measured over time. Lovenheim (2011) shows that these self-reported housing values match up closely with national trends in housing prices, suggesting that self-reports contain little systematic bias. Home ownership status is defined as owning a home in the survey year.[9] For renters, the market housing price is the mean housing price in their MSA and survey year from all home owners in the sample. We also control for women's age categories, women's education attainment levels, real family income, marital status, and the number of other children in the home.[10] Women who have moved in the prior four years are dropped from the sample.

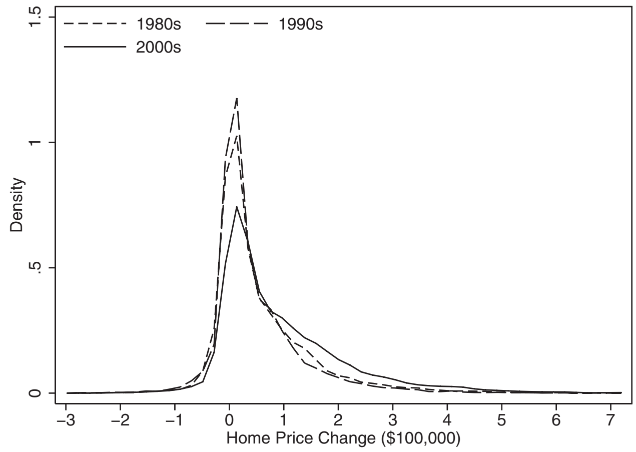
Table 1 contains summary statistics of the PSID data we use, separately for home owners and renters. The table shows that renters have higher fertility rates than home owners, at 5.4% versus 5.0%. Renters also tend to live in areas with higher housing prices, are less likely to be married, are younger, and are less educated. The mean home value among home owners is about $159,000, but the variance of this variable is large. Similarly, the average home price increase over two years is about $35,000, and over four years it is over $60,000; both measures exhibit a large amount of variation in the data, with standard deviations significantly larger than the means.

While there are both increases and decreases in home prices, there are few decreases relative to increases. Over

[9] A possible objection to measuring home ownership in the survey year is that it is endogenous. When we define home ownership with a two- or four-year lag, our results are quantitatively and qualitatively similar. These results are available from the authors on request.

[10] We cannot determine in our data whether a child living in a home with a woman is that woman's child. We use the number of other children who were not just born as a proxy for the number of children to whom a woman has given birth. Given the data limitations, this method is a reasonable one for controlling for the fact that the fertility hazard declines with the number of existing children.

FIGURE 3.—THE DISTRIBUTION OF FOUR-YEAR HOME PRICE CHANGES AMONG FEMALE HOME OWNERS AGES 25–44, BY DECADE



Source: Panel Study of Income Dynamics sample of female home owners, ages 25–44.

a four-year period, only 16.8% of the home value changes are negative because the large housing market bust occurred largely after the 1985–2007 time period we examine. There is a great deal of variation in the home value changes that increased over time. Figure 3 shows the distribution of four-year home price changes among home owners in our sample by decade. In the 2000s, the distribution shifted outward above 0. That this shift did not occur in the negative range demonstrates that our data do not capture the majority of the housing bust. It also is important to note that home price increases, particularly during the housing boom, were not limited to high-housing-price states or cities and were not limited to wealthy home owners. Rather, many historically lower-price cities and many lower-income individuals in those cities experienced large wealth increases from the housing boom.[11]

IV. Estimation Strategy

We estimate linear probability models of the following form for home owners on the PSID data described in the previous section from 1985 to 2007:[12]

$$birth_{ist} = \beta_0 + \beta_1 \Delta P_{ist} + \gamma X_{ist} + \theta_s + \phi_t + \eta_{ist}, \qquad (2)$$

where i indexes women, s indexes state or MSA (depending on the specification), and t indexes survey year. The variable P is home price, and X is the set of observable characteristics shown in table 1, as well as the state-by-year average unemployment rate and log real income per capita. The θ_s are state or MSA fixed effects, and ϕ_t are year fixed effects. The main variable of interest in equation (2) is ΔP, the individual home price change over the past two or four years among home owners. We focus on recent individual-level changes because contemporaneous home price levels may be only loosely related to household resources: families can own an expensive home without having any equity in it. However, individual home price changes are capitalized into housing wealth, which makes home price changes a more appropriate measure of wealth than the home price level. The coefficient of central importance in equation (2), β_1, thus shows how the likelihood of having a child in the previous year is associated with recent home price changes. To demonstrate the importance of using home price changes rather than levels, we show that fertility does not respond to the contemporaneous value of homes.

The identification assumption underlying equation (2) is that housing price changes are conditionally exogenous to the fertility decision. In other words, but for the fact that housing prices increase household wealth, home price changes and fertility should be uncorrelated conditional on the observables in the model. There are two main threats to this assumption. The first is a positive correlation between housing prices and local macroeconomic conditions. If fertility responds positively to macroeconomic variation, our housing price change measures may be picking up this relationship rather than identifying the effect of housing wealth changes on fertility decisions.[13] To guard against this possibility, we control for the state-average unemployment rate and real income per capita, direct measures of state-level macroeconomic conditions. Furthermore, we estimate our model for renters using average home price changes in their locality. Because renters experience local macroeconomic changes similar to those of home owners but not any commensurate wealth changes due to housing market variation, these estimates lend insight into the extent of any bias in our estimates that is driven by unobserved macroeconomic factors.

We also estimate a model using state-by-year fixed effects, which will control for all unobserved factors common within state and year. With such fixed effects, β_1 is identified off of housing price growth differences among home owners within a state and year. While it still is possible for these within-state and year differences to be driven by economic shocks, the local dynamics of housing price changes within a state is more likely to be driven by exogenous factors, such as local supply constraints, that are less prone to bias from macroeconomic shocks than is within-MSA variation over time (Gyourko, Mayer, & Sinai, 2006; Glaeser, Gyourko, & Saks, 2005). Most notably, this method controls for uniform state-year-level economic shocks. The estimates using state-by-year fixed effects (reported in table A1 of the table appendix) are very similar to the baseline models discussed below, which is

[11] For example, in the 2001–2007 PSID samples, the average four-year increase that home owners experience in the lower half of the income distribution is 34%, and it is 41% among the top half of the income distribution.

[12] Given the low likelihood of birth in each year, it is not clear that a linear model is appropriate. However, marginal effects from a logit model yield similar results. These estimates are available from the authors on request, but we report linear probability model coefficients due to their ease of interpretation.

[13] We are not aware of any convincing evidence that such a relationship exists. In fact, aggregate trends of macroeconomic conditions and fertility would suggest that they are negatively correlated.

suggestive that macroeconomic shocks are not confounding our estimates.

The second potential threat to identification is the selection of households across states, across MSAs within states, or across neighborhoods within MSAs. If women who are planning to have children purchase homes in places that are most likely to experience high housing price growth in the near future (perhaps because of a good school district, city parks, or other amenities), our estimates will be biased upward. In order to address this problem, we use successively more restrictive housing price growth variation to estimate equation (2). First, we employ state fixed effects, which allow housing price growth across time within states as well as cross-sectionally across MSAs and neighborhoods within an MSA. We then include MSA fixed effects using the restricted-access geocodes from the PSID. This specification allows differential growth rates across time within MSAs and across neighborhoods cross-sectionally within an MSA. Finally, we use simulated housing price growth. Household i's simulated home price in MSA s at time t conditional on its $t-4$ home price is

$$\hat{P}_{ist} = P_{is,t-4} * \frac{HPI_{st}}{HPI_{s,t-4}}. \qquad (3)$$

This simulated home price forces all growth between $t-4$ and t to be due to MSA-level housing price growth, which we calculate using the MSA-level housing price index (HPI) provided by the Federal Housing Finance Agency. We calculate the simulated housing price growth for two and four years, defined as $\hat{P}_{ist} - P_{ist-2}$ and $\hat{P}_{ist} - P_{ist-4}$, respectively.

Conditional on MSA fixed effects, these simulated home price growth measures allow only within-MSA variation in housing price growth rates over time. In this specification, our estimates of β_1 will be biased upward only if household selection patterns change over time across MSAs in such a way that families that are more likely to have a child in the near future begin moving disproportionately into MSAs where the housing price growth will be the highest. We believe such selection changes are implausible, and we know of no evidence suggesting that migration patterns changed in this way over the past fifteen to twenty years. To guard against selection of this type, we construct the simulated housing price growth measure using the MSA of the original 1968 respondent rather than the actual MSA of residence.[14] By not allowing any migration, estimates from this specification will not be biased by selective migration.

Although sampling in the PSID is done at the household level, there is a strong within-MSA component to home price change variation. Due to the strong common geographic component to home price variation, it is likely that errors are not independent within cities. We therefore cluster our standard errors at the MSA level throughout. For respondents who do not live in an MSA, we create a separate non-MSA cluster for each state. This method allows arbitrary correlation of the errors within each state for rural respondents and within each MSA for those living in an identifiable city. Standard errors that are clustered at the household level are similar, however.

V. Results

A. *Results for Home Owners and Renters*

Results from estimation of equation (2) are shown in table 2. Each column of the table presents results from a separate regression, and all estimates include the full set of controls shown in table 1 as well as state macroeconomic controls. The first three columns of panel A show our estimates for home owners when state fixed effects are included in the model. In the first column, there is a weak association between housing price levels and the likelihood of giving birth. This finding could be either because there is little relationship between housing wealth and fertility or because housing prices are a poor measure of housing wealth. In columns (2) and (3), respectively, we show the effect of a two-year and four-year change in housing values among home owners on the likelihood of giving birth. When the two-year change is used, a \$100,000 change in home prices leads to a 0.0089 percentage point, or 17.8%, change in the likelihood of having a child in the past year. Using the four-year change, a \$100,000 change in home prices leads to a 0.0082 percentage point, or 16.4%, change in the likelihood of having a child in the past year. These estimates are significantly different from 0 at the 5% level.

These marginal effects of a \$100,000 increase in home value may appear small, but due to the large increases in home prices in the early 2000s and the low underlying birth probability, they translate into sizable responses. During the housing boom from 1999 to 2005, the average two-year home price increase among home owners was \$48,024, and the average four-year home price increase was \$77,911. These increases translate into total increases in the likelihood of having a child among home owners of 0.43 and 0.64 of a percentage point, respectively. Compared to the baseline fertility rate of 5.0% for home owner women in our sample, housing wealth increases in the early 2000s increased fertility by 8.6% (0.43/5.0) for the two-year change and 12.8% (0.64/5.0) for the four-year change. Thus, the recent variation in housing prices has been large enough to generate economically meaningful changes in fertility among home owners. The average home equity in our sample is \$74,086, which implies that the housing wealth elasticity of fertility is 0.13.[15]

The assumption underlying the identification of β_1 in equation (2) is that households with higher underlying fertility

[14] For the 1990–1992 PSID resample and the 1997 immigrant sample, we use the MSA when the individual was age 20.

[15] The average home equity in our sample is \$74,086. A \$100,000 increase in home price is a 134.98% increase in housing wealth. Combined with the 16.4% increase in the probability of having a children from a \$100,000 increase in home price, this implies a housing wealth elasticity of 0.13. Alternatively, we can compute the total wealth elasticity. The average home owner age 25 to 44 has about \$120,000 in total wealth. Therefore, the elasticity with respect to total wealth is about 0.20.

TABLE 2.—LINEAR PROBABILITY MODEL ESTIMATES OF THE EFFECT OF HOUSING PRICES ON BIRTH PROBABILITY

Independent Variable	(1)	(2)	(3)	(4)	(5)	(6)
A: Home Owners						
Home value ($100,000)	0.0016			0.0020		
	(0.0011)			(0.0013)		
2-year home value change ($100,000)		0.0089**			0.0088**	
		(0.0018)			(0.0019)	
4-year home value change ($100,000)			0.0082**			0.0085**
			(0.0015)			(0.0002)
Real family income ($100,000)	−0.0009	−0.0004	−0.0017	−0.0009	−0.0001	−0.0013
	(0.0019)	(0.0018)	(0.0018)	(0.0020)	(0.0020)	(0.0020)
MSA fixed effects?	No	No	No	Yes	Yes	Yes
R^2	0.055	0.056	0.056	0.069	0.069	0.070
B: Renters						
Home value ($100,000)	−0.0047			0.0035		
	(0.0048)			(0.0030)		
2-year home value change ($100,000)		−0.0015			0.0001	
		(0.0062)			(0.0045)	
4-year home value change ($100,000)			−0.0001			0.0019
			(0.0052)			(0.0031)
Real family income ($100,000)	−0.0093*	−0.0079	−0.0084	−0.0049	−0.0041	−0.0046
	(0.0050)	(0.0054)	(0.0057)	(0.0055)	(0.0060)	(0.0063)
MSA fixed effects?	No	No	No	Yes	Yes	Yes
R^2	0.039	0.040	0.041	0.045	0.046	0.049

Dependent variable: Dummy = 1 if gave birth in the previous year. Authors' estimation of equation (2). In columns 1–3, respondents not in an MSA are assigned to a non-MSA state-specific cluster. Estimates in columns 4–6 use only respondents who live in an identifiable MSA at the time of the interview. All estimates include state and year fixed effects, age group dummies (with 25–29 as the excluded category), educational attainment dummies (with no high school diploma as the excluded category), and controls for marital status, the number of other children in the household, state-by-year unemployment rates, and state-by-year real income per capita. For the renter sample in panel B, housing price measures are calculated using home owners within each MSA and year as described in the text. Standard errors clustered at the MSA level are in parentheses. Significant at **5% and *10%.

Source: 1985–2007 Panel Study of Income Dynamics, women ages 25–44; those who moved in the prior four years are excluded.

rates are not sorting into regions in which housing prices are growing the fastest.[16] In columns 4 to 6 of table 2, we include MSA fixed effects that control for the systematic differences among households across MSAs within states in underlying fertility rates. Although these fixed effects significantly reduce the housing price change variation, the estimates are virtually identical to those using state fixed effects.

In every column of panel A of table 2, the income coefficient is negative and insignificant. The magnitudes range from −0.0001 to −0.0017 for a $100,0000 change in family income. While these estimates can be interpreted as indicating a negative effect of family income on fertility, we urge caution in such an interpretation because we lack an instrument to generate exogenous income variation in our sample. These negative coefficients likely are driven by many of the same biases that drive the negative cross-sectional correlation between fertility and income at the aggregate level. We include income in our models as a control variable that provides an important measure of each woman's economic circumstances, but the coefficient on income likely does not identify a causal effect. Importantly, all of our estimates are robust to excluding family income from our models.

One test of whether the effects we are estimating can be attributed to wealth rather than to an MSA-level shock that is correlated with both housing prices and fertility or to a substitution effect is to estimate our model for renters. For renters, the association between housing price and fertility is likely negative (or at least nonpositive). Rising home prices do not provide a wealth increase to renters but may increase rental prices and cause a substitution effect. We therefore estimate a version of equation (2) using MSA-by-year average home prices for home owners as the measure of housing price for renters. Panel B of table 2 presents these estimates, with specifications that are the same as those for the corresponding columns of panel A.

The negative coefficients on home price changes in columns 2 and 3 are suggestive of a small substitution effect; however, the estimates in columns 5 and 6 are positive, suggesting no substitution effect. Regardless, in no column is the home price estimate statistically significantly different from 0 at even the 10% level. While these results provide supporting evidence for the causal interpretation of our estimates among home owners and suggest that the negative bias stemming from income effects in our estimates are small, we caution against direct comparisons because as table 1 shows, renters are quite different from home owners along several dimensions. Our results using the PSID data among home owners unequivocally reject a negative relationship between housing wealth and fertility and strongly suggest that those who experience housing wealth increases are more likely to have children, thus suggesting that any substitution effect is swamped by the wealth effect of home price increases.

Even with MSA fixed effects, it still is possible that our housing price growth estimates are biased due to selection of households with different fertility patterns into different neighborhoods with systematically different housing price growth within an MSA. To account for this possibility, we estimate equation (2) using simulated price changes within

[16] Note that the majority of estimates of the effect of household income on fertility suggest such selection will bias us against finding positive results because higher-income families have fewer children.

TABLE 3.—LINEAR PROBABILITY MODEL ESTIMATES OF THE EFFECT OF HOUSING PRICES ON BIRTH PROBABILITY FOR HOME OWNERS USING SIMULATED MSA-LEVEL HOUSING PRICE CHANGES

	Current MSA		Original MSA	
Independent Variable	(1)	(2)	(3)	(4)
2-year home value change ($100,000)	0.0073**		0.0075**	
	(0.0016)		(0.0016)	
4-year home value change ($100,000)		0.0057**		0.0055**
		(0.0012)		(0.0012)
Real family income ($100,000)	−0.0003	−0.0007	−0.0002	−0.0001
	(0.0020)	(0.0020)	(0.0020)	(0.0020)
R^2	0.068	0.069	0.067	0.068

Dependent variable: Dummy=1 If gave birth in the previous year. Authors' estimation of equation (2) using simulated housing price changes that use only MSA-level variation in housing price growth rates. Columns 1 and 2 use the simulated housing price change for the MSA the individual lived in for at least the past four years. Columns 3 and 4 use the simulated housing price change for the MSA of the original 1968 respondent or the MSA when the individual was age 20 for those added to the PSID in the 1997 immigrant sample or the 1990–1992 resample. All estimates include MSA and year fixed effects, age group dummies (with 25–29 as the excluded category), educational attainment dummies (with no high school diploma as the excluded category), and controls for marital status, the number of other children in the household, state-by-year unemployment rates, and state-by-year real income per capita. The estimation sample includes only respondents who live in an identifiable MSA at the time of the interview and own a home. Standard errors clustered at the MSA level are in parentheses. Significant at **5% and *10%.

Source: 1985–2007 Panel Study of Income Dynamics, women ages 25–44, those who moved in the prior four years are excluded.

current and original MSAs, a method that restricts housing price growth rates to be the same in each year in each MSA.[17] table 3 shows the results using these price changes. The estimates are somewhat smaller than those in table 2: a $100,000 increase in home prices in the previous two years is associated with a 0.0073 percentage point increase in the likelihood of giving birth, and a $100,000 increase in home prices in the previous four years is associated with a 0.0057 percentage point increase in the likelihood of giving birth. Though these estimates are smaller, they still are statistically different from 0 at the 5% level and indicate a sizable effect of housing wealth on fertility. The smaller magnitude suggests that selection of women with higher fertility to neighborhoods with higher housing price growth may have caused a small positive bias in the results shown in table 2. The implied housing wealth elasticity using the four-year home value change is 0.08 as compared to an elasticity of 0.13 implied by using the actual home price change. Defining the simulated price growth by using the original rather than the current MSA has no effect on the results and indicates that selective migration across MSAs is not causing any bias in our estimates.

TABLE 4.—LINEAR PROBABILITY MODEL ESTIMATES OF THE EFFECT OF FOUR-YEAR HOUSING PRICE CHANGE ON BIRTH PROBABILITY FOR HOME OWNERS, BY AGE

Four-Year Home Price Change Interacted with:	Estimated Effect of $100,000 Increase (1)	Mean Fertility Rate (2)	Estimated % Change in Birth Probability (3)
Age group			
15–19	−0.0059**	0.0327	−18.04%
	(0.0016)		
20–24	−0.0049	0.0740	−6.62
	(0.0037)		
25–29	0.0175**	0.1172	14.93
	(0.0054)		
30–34	0.0144**	0.0865	16.65
	(0.0040)		
35–39	0.0081**	0.0263	30.80
	(0.0020)		
40–44	0.0042**	0.0056	75.00
	(0.0014)		

Authors' estimation of equation (2) with interaction terms for each age group in the regression. The specification includes MSA and year fixed effects, age group dummies, educational attainment dummies, and controls for marital status, the number of other children in the household, state-by-year unemployment rates, and state-by-year real income per capita. The estimation sample includes only respondents who live in an identifiable MSA at the time of the interview who own a home. Standard errors clustered at the MSA level are in parentheses. Significant at **5% and *10%.

Source: 1985–2007 Panel Study of Income Dynamics, women ages 15–44; those who moved in the prior four years are excluded.

B. *Is This a Fertility Timing or a Total Fertility Effect?*

The estimates thus far suggest that families experiencing higher home price growth are more likely to have a child in a given year, which could indicate changes in the timing of fertility or in total fertility over the life course. In table 4, we estimate the effect of home price changes on total fertility by interacting four-year home price changes with age group indicators. For this specification, we have included women age 15 to 24 along with the women age 25 to 44 used in all other specifications. For these young women, especially those in the 15–19 age group, home ownership likely means that the woman lives with her parents, who own the home. The estimated effect of a $100,000 increase in home value (to the parents) for these 15- to 19-year-old women is an 18% reduction in the likelihood of giving birth. Though the negative relationship between family income and fertility of young women is well documented, this is the first evidence of which we are aware indicating a causal negative relationship between family wealth and fertility for teenage women.

The results reported in table 4 also indicate that the response among 25- to 44-year-old-women is not only being driven by younger women. In fact, women in all age groups over 25 respond positively to home price changes, with the largest percentage change in birth probability occurring for the 35- to 44-year-old-women. Taken together, these estimates suggest that home price changes do not solely induce a

[17] This method significantly restricts the variation in home price changes across households within the same MSA in a given year. However, at a constant growth rate, the level of growth will be higher for more expensive homes. In results not reported, we have controlled for lagged home price levels, and the estimates on the simulated home price change variables are virtually unchanged. The estimates are available from the authors on request and suggest that the identifying variation in these models is coming from home price growth rate differences, which do not vary cross-sectionally within MSAs by construction.

TABLE 5.—THE EFFECT OF CURRENT AND LAGGED HOUSING PRICE CHANGES ON FERTILITY

Independent Variable	Dependent Variable: Dummy=1 If Gave Birth in Previous Year (1)	Dependent Variable: Number of Children Born over Prior Four Years (2)
4-year change	0.0079**	0.0286**
	(0.0024)	(0.0110)
Lagged 4-year change	0.0009	−0.0044
	(0.0018)	(0.0102)
R^2	0.072	0.253

Authors' estimation of equation (2) using both four-year housing price change ($P_t - P_{t-4}$) and lagged four-year housing price change ($P_{t-4} - P_{t-8}$). All estimates include MSA and year fixed effects, age group dummies, educational attainment dummies, and controls for marital status, the number of other children in the household, state-by-year unemployment rates, and state-by-year real income per capita. The estimation sample includes only respondents who live in an identifiable MSA at the time of the interview and who own a home. Standard errors clustered at the MSA level are in parentheses. Significant at **5% and *10%.

Source: 1985–2007 Panel Study of Income Dynamics, women ages 25–44; those who moved in the prior four years are excluded.

TABLE 6.—LINEAR PROBABILITY MODEL ESTIMATES OF THE EFFECT OF FOUR-YEAR HOUSING PRICE CHANGE ON BIRTH PROBABILITY FOR HOME OWNERS, BY NUMBER OF CHILDREN, FAMILY INCOME, AND DECADE

Four-Year Home Price Change Interacted with:	Estimated Effect of $100,000 Increase (1)	Mean Fertility Rate (2)	Estimated % Change in Birth Probability (3)
Number of children			
0 children	−0.0007	0.0428	−1.64%
	(0.0025)		
1 child	0.0305**	0.0727	41.95
	(0.0048)		
2 children	0.0041*	0.0448	9.15
	(0.0024)		
3+ children	0.0056*	0.0494	11.34
	(0.0032)		
Family income			
Top quartile	0.0103**	0.0543	18.97
	(0.0023)		
Third quartile	0.0105**	0.0605	17.36
	(0.0049)		
Second quartile	0.0076**	0.0515	14.76
	(0.0038)		
Bottom quartile	−0.0018	0.0427	−4.22
	(0.0034)		
Decade			
1985–1989	0.0046*	0.0618	7.44
	(0.0032)		
1990–1999	0.0098**	0.0517	18.96
	(0.0027)		
2001–2007	0.0101**	0.0415	24.34
	(0.0027)		

Authors' estimation of equation 2 with interaction terms for specified observable characteristics. Each panel is a separate regression, and all specifications include MSA and year fixed effects, age group dummies, educational attainment dummies, and controls for marital status, the number of other children in the household, state-by-year unemployment rates, and state-by-year real income per capita. The estimation sample includes only respondents who live in an identifiable MSA at the time of the interview and own a home. Standard errors clustered at the MSA-level are in parentheses. Significant at **5% and *10%.

Source: 1985–2007 Panel Study of Income Dynamics, women ages 25–44; those who moved in the prior four years are excluded.

change in the timing of fertility but actually cause an increase in total family size. By multiplying each of the coefficients by 5, one can calculate the effect on total fertility. This calculation indicates that a $100,000 increase in housing wealth increases the total number of children born to women ages 25 to 44 by 0.22, which is 18.8% of the underlying total fertility rate for this population. Considering all women of childbearing age, total fertility increases by 0.17, a 9.8% increase relative to baseline.

One potential criticism of our calculation of total fertility effects is that the calculation assumes that conditional on the four-year home price change, fertility decisions are not responding to more lagged changes. For example, if a women experiences a home price increase when she is 29, this may reduce the likelihood she gives birth when she is 35. Such responses to lagged home price changes will cause us to overstate the total fertility effect using the estimates from table 4. Thus, in table 5, we present results from estimates of equation (2) with both the four-year home value change ($P_t - P_{t-4}$) and the four-year lagged four-year home value change ($P_{t-4} - P_{t-8}$). If the increased probability of birth was primarily a timing shift, we would expect to see a decrease in the probability of birth in the years after the response. Column 1 of table 5 indicates that this is not the case. The lagged four-year change has no effect on the probability of birth, while the parameter estimate for the current four-year change is similar to those from our baseline specifications. In column 2, we use the number of children born over the prior four years, rather than a birth indicator, as the dependent variable. Here, the parameter estimate on the lagged four-year change is negative but small and not statistically significant. The evidence in tables 4 and 5 is suggestive that the fertility response we find is not simply due to birth timing but indicates an increase in completed fertility.

C. *Response Heterogeneity*

The average effects of housing wealth on fertility point to children being normal goods, with the added wealth from home price increases driving higher fertility. However, the fertility response is likely quite heterogeneous. To examine several sources of heterogeneity, we interact the four-year home price change with indicators for various characteristics. These results are presented in table 6.

The top of table 6 reports results from interacting the four-year home price change with the number of children already born to the woman. We find that there is no response by women who do not already have a child. This finding helps alleviate concern about selection into home ownership by women who plan to have a child being correlated with the housing price change. If these women were more likely to buy a home in an area with higher future housing price growth, for example, it could create a positive bias in our estimates. The results indicate that women with one child are more responsive than women with two or more children. Although smaller, the estimated percentage change is still quite large for women with two children and for those with three or more children.

Estimates by family income are reported in the second panel of table 6. We find little evidence of heterogeneity in the response across families in the top three-quarters of the income distribution, but the fertility decisions of families in the lowest income quartile are not responsive to changes in housing wealth. That we find no fertility response for these low-income women suggests that they either have a lower wealth elasticity of fertility or a different reaction to an increase in housing wealth than high-income women. While beyond the scope of this analysis, future work focusing on understanding the underlying causes of this heterogeneity would be of value.

The third panel of table 6 examines the effect of housing wealth on fertility by decade in order to determine whether our estimates are driven, at least in part, by the increased availability of home equity loans in the 1990s and 2000s. Consistent with this hypothesis, the estimated response is lower in the 1980s than in the subsequent decades and has grown steadily over time, from a 7.4% effect in the 1980s to a 24.3% effect in the 2000s for each $100,000 increase in home price. These estimates are suggestive that the decline in housing prices from the recent housing bust, as well as the reduced access to home equity loans and lines of credit, may have had a significant impact on household fertility decisions. Because our data do not cover the housing bust period, we cannot examine such effects directly, and using these estimates to predict the effect of the housing market decline on fertility requires a symmetry assumption. Using the 16.8% of our sample that experiences price declines, we found some evidence that the response is not symmetric: although the effect of a decrease in home value is not statistically different from 0, the standard errors are quite large. In addition, the home price declines in our sample come predominantly from earlier periods, when reductions in home prices were not accompanied by large reductions in the liquidity of housing wealth. Our estimates indicate that the combination of high home price growth and liquid housing wealth during the recent housing boom led to fertility becoming more responsive to housing market variation. More work examining the effect of the housing bust on fertility and on household behavior more generally is needed in the future when data from the period of the housing bust become available.

VI. Conclusion

We use housing market variation to estimate the fertility response to a change in housing wealth using individual-level data from the PSID. We find a positive and significant effect of both two-year and four-year housing price growth on the likelihood that a woman has a child in the preceding year. Though the marginal effect of a $100,000 increase appears small, at about 0.85 of a percentage point, we argue that this partial effect is economically significant given the 0.05 baseline probability of birth and the large variation in housing prices experienced over the past decade. Furthermore, our estimates are robust to using successively more restrictive housing price growth measures, which suggests that selection of households across MSAs or across neighborhoods within MSAs is not driving our results. We also find no evidence of a commensurate fertility response among renters, suggesting that our results are not driven by location-specific economic conditions or other location-specific time-varying factors.

Our results are consistent with a small but growing body of literature that calls into question the conventional wisdom that fertility and family resources are negatively linked. Our focus on wealth in general and on housing wealth in particular is unique in this literature, and it allows us to generate estimates of the effect of family resources on fertility without using wage shock measures that could affect the relative trade-off of home versus market work. That we find similar results using a new source of identification provides support to the other papers in this literature.

Using the age distribution of women in our sample, we show that home price changes affect total fertility rather than only the timing of fertility. Our estimates indicate that a $100,000 increase in housing wealth increases total fertility by 0.22 children, an 18.8% increase. We also provide evidence that the negative correlation between family resources and the fertility of teenage women is at least partially causal: an increase in housing wealth seems to cause a reduction in teenage births.

Finally, that we estimate heterogeneous treatment effects by the number of existing children, family income, and decade is unique to this literature. In particular, we show that the decision to have the first child is unaffected by home price growth and that the estimated response is due to those women in the top three-quarters of the family income distribution. Our results further indicate that the responsiveness of fertility to home price changes has grown over time with the liquidity of housing wealth, which has important policy implications given the severe reduction in the liquidity of this asset during the housing boom.

This paper also contributes to the literature on the effects of housing market fluctuations by demonstrating that fertility choices are among the set of behaviors that are influenced by changes in the housing market. Other work that has shown housing market effects includes Davidoff (2010), who shows that the demand for long-term care insurance is negatively affected by housing wealth. Lovenheim (2011) shows that housing market changes affect college attendance, and Lovenheim and Reynolds (2011) present evidence that this variation has an impact on college choice and the likelihood of graduation as well. Many authors, including Campbell and Cocco (2007), Case, Quigley, and Shiller (2005), Hurst and Stafford (2004), and Lehnert (2003), have found that housing wealth affects consumption.[18] Our results add to this literature by showing that the recent severe declines in the housing market also may have important fertility consequences, although our results indicating that effects are not symmetric with respect to home price increases and decreases points to the need for more direct study in the future of how the housing decline impacted fertility decisions.

[18] Attanasio et al. (2009) take issue with this interpretation of this literature, however, and argue this relationship is incidental.

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TABLE APPENDIX

TABLE A-1.—LINEAR PROBABILITY MODEL ESTIMATES OF THE EFFECT OF HOUSING PRICES ON BIRTH PROBABILITY FOR HOME OWNERS WITH STATE-BY-YEAR FIXED EFFECTS

Independent Variable	(1)	(2)	(3)
Home value (\$100,000)	0.0012		
	(0.0011)		
2-year home value change (\$100,000)		0.0087**	
		(0.0019)	
4-year home value change (\$100,000)			0.0085**
			(0.0015)
Real family income (\$100,000)	−0.0002	−0.0001	−0.0014
	(0.0020)	(0.0019)	(0.0019)
R^2	0.082	0.083	0.083

Dependent variable: Dummy=1 If give birth in the previous year. Authors' estimation of equation 2. All estimates include state-by-year fixed effects, age group dummies (with 25–29 as the excluded category), educational attainment dummies (with no high school diploma as the excluded category), and controls for marital status and the number of other children in the household. Standard errors clustered at the MSA level are in parentheses. Respondents not in an MSA are assigned to a non-MSA state-specific cluster. Significant at **5% and *10%.

Source: 1985–2007 Panel Study of Income Dynamics, women ages 25–44; those who moved in the prior four years are excluded.