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The Joint Dynamics of Family Housing and Childbearing Decisions

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November 12, 2020

Abstract

Housing price growth is now and then attributed as one reason of fertility decline in cities. As the cost of housing keeps climbing over years, it is likely the increasing financial burden not only bars new home buyers from entering homeownership but also has an impact on their family plan to raise children. The direction of the impact is however unclear. By formulating the family behavior of housing and childbearing as a joint decision-making process, we investigate into the effect of local housing price variation on the both behaviors simultaneously for non-homeowning women in the United States. We estimate a multinomial logit model of the interaction of the two binary choices, entering homeownership and giving a birth, using the family data from the Panel Study of Income Dynamics between 1985 and 2015 and the corresponding MSA-level house price data imputed from the Federal Housing Financial Agency and the Census. The results show that, high house price level strongly discourages the probability of entering homeownership, while it has a very mild net positive relation with the likelihood of childbearing for non-homeowning women. In areas with high house price, families are more likely to have a new baby before buying a home mostly

because of the substantial drop of the probability of entering homeownership and giving a birth in tandem in one or two years. Though the net effect on childbearing is small, high house price would nonetheless raise the chance of parenting without homeownership. On the other hand, the effect of house price change, regardless the price level, is hardly found.

Note

Some of the data used in this analysis are derived from Sensitive Data Files of the Panel Study of Income Dynamics, obtained under special contractual arrangements designed to protect the anonymity of respondents. These data are not available from the authors. Persons interested in obtaining PSID Sensitive Data Files should contact through the Internet at PSIDHelp@isr.umich.edu.

1 Introduction

The decision to buy a home and to rear children are two major and interrelated decisions of a family. On the one hand, home and children are somewhat complementary to achieve a paradigm family in most cultures. Especially in where nuclear family is the prevailing family ideal, owning a home in general gives a family more and usually better living space to raise children (Haurin et al., 2002; Kulu and Vikat, 2007), and in some countries it even serves as a qualification of formal family formation (Wei et al., 2012). Moreover, the desire to have a bigger family usually calls for a higher demand for housing space (Mulder, 2006b; Kulu and Steele, 2013). On the other hand, both homeownership and parenting are constrained by a single family budget, and home price variations in the market can theoretically affect the fertility choices of families who seek for them both due to the economic reason. For potential home buyers (non-homeowners), an increase in housing costs would crowd out the expense for child-rearing and force them to postpone or even give up one of the two goals. For homeowners, although the added wealth from property appreciation would ease the family financial burden and allow them to raise more children (Clark and Ferrer, 2019; Cloyne et al., 2019), their ability to move to a larger home could be limited. This interplay of homeownership, childbearing, and housing market implies a complicated relation between house price and fertility. The impact of growing housing cost is unlikely to be confined in the tightened budget for consumption. The welfare loss of delayed family building can also be a consequence, especially for non-owners.

Now and again, the issue of climbing house price enters the center of public policy debates such as on public housing and inequality, city gentrification and redevelopment, home mortgage market regulations, and the policy response to low fertility rates. The association between housing and fertility decisions has attracted extensive attentions across multiple fields of the social sciences in the past decade. It is however difficult to disentangle the interaction between homeownership and childbearing in analysis because of their interweaving influences on each other's decision. Recent empirical literature has been investigating the effect of house price on fertility decision mostly by restricting the research sample to either the homeowners or the renters to explore the one-to-one causality between house price and childbearing choice to prevent the confounding factor of homeownership. Notwithstanding their findings, the underlying mechanism of the relation between house price and childbearing decision is yet to be touched. House price variation has a more direct impact on home buying decision. As homeownership and childbearing are correlated, home buying decision can be deeply involved in the childbearing decision, and *vice versa* (Clark and Withers, 2009). Focusing on the influences on childbearing choice only tends to overlook the variation in homeownership transition and its subsequent effect on childbearing for either group. This omission may downplay the negative effect of growing house prices on the well-being of non-homeowners.

We are interested in how local housing market affects the dynamics of the family decision of homeownership transition and childbearing of potential home buyers. Not only we revisit the question about the impact of high local house price level on childbearing decision, we also investigate how the variation in price makes an impact on family life course plan of both housing and parenting. Does high house price shift the order of the transitions? Does the booming housing market have the same effect? These are the questions in our mind. To answer them, we estimate a multinomial logit (MNL) model of the crossing of the two binary choices, becoming homeowner and giving a birth, on local median house price level and change for women living in the U.S. between year 1985 and 2015. The data mainly comes from the Panel Study of Income Dynamics (PSID) and the corresponding MSA-level house price data imputed from the Federal Housing Financial Agency (FHFA) and the Census.

The results draw a pattern of the impact of the housing market on the behavior of non-homeowning women in the U.S. High house price level strongly discourages the probability of entering homeownership, while it has a very mild net positive relation with the likelihood of childbearing for non-homeowning women. In areas with high house price, families are more likely to have a new baby before buying a home mostly because of the substantial drop of the probability of entering homeownership and giving a birth alongside in one or two years (hereafter called as the "doing both" move). The effect of house price change, regardless the price level, is hardly found. Only a negative effect of four-year price change on doing both is observed from the analysis. Overall, house price level, instead of short-term variation in price, has more prominent impact on the family choice on homeownership and childbearing. An analysis on the interaction effect also shows the effect scales differ by family income level, partnership, and race. Though the net effect on childbearing is small, high house price would nonetheless raise the chance of parenting without homeownership.

Unlike Beckerian static theoretical models, both the two decisions are dynamic over time. If a family is heading for both homeownership and new child while the credit constraint is binding, it must postpone its plan, save money and wait for better chance, and now and again adjust to the varying costs. Facing a growing housing price, a renting family would either delay having a child to accumulate the mortgage down payment first or have a child right away and postpone the plan to enter homeownership. The relative costs and preferences on homeownership and children are critical in making the decision. The recent price trend can also shape the expected price for the future and alter the decision if the forward-looking perspective is considered.

The knowledge on the association of housing and fertility, especially to the behavior of newly formed families, is important to the assessment of relevant housing policies. Its efficacy is closely tied with family responses to the change in their financial budget. Our results suggest a small side effect on childbearing could accompany with policies targeting on the homeownership affordability. A better understanding on the dynamics of family decision will help policy maker to predict both the direct and indirect outcome of those policies on either homeownership rate or birth rate more accurately, and the evaluation on relevant policies will be more conclusive and precise.

2 Background

Financial consideration is one major material restriction for family to have more and better care for children in industrialized countries. Since housing is typically the largest part of living expense and often the largest store of wealth for families, it is natural to wonder how local house prices could affect family size, especially as many urban areas has experienced property appreciation recently. A growing empirical literature in recent years investigate into this relationship around the world thanks to the increasing data accessibility of housing markets and demography, but the findings diverge with different methodologies and data source. Some time series (e.g. Yi and Zhang (2010) on Hong Kong) and cross-sectional regression (e.g. Simon and Tamura (2009) on the U.S.) studies conclude that high house price has a negative impact on local fertility rate. Others cast doubts on such conclusion and instead argue for a positive relation through the wealth channel. For instance, Feyrer et al. (2008) argues that the U.S. data shows a positive correlation, though weak, between housing prices and fertility rates.

Among them, a few studies isolate the wealth effect of property appreciation by controlling homeownership. Dettling and Kearney examine the relationships of the lagged house price index and homeownership rate on the MSA-level fertility rates and finds that the positive home price effect is greater in areas with higher homeownership rate, and the prediction implies that the overall relation between home price and fertility is slightly positive in the United States. Lovenheim and Mumford employs a panel dataset of the U.S. households between 1985 and 2005 from the PSID data to show a positive effect on homeowners but find no significant repercussion for renters. Clark and Ferrer find a similar result for homeowners and non-owners using a Canadian longitudinal data, and lately Daysal et al. (2020) reports an empirical finding that the observed effect for homeowners in Denmark has almost the same scale with that observed in the U.S. By and large, the leading evidence suggests that house price appreciation generates a dominating positive wealth effect on childbearing for homeowners but only a weak negative to none price effect for non-homeowners (renters) on the likelihood of childbearing for families in general.

The association of housing and childbearing is yet to be fully explained. Though these empirical works directly estimate the impacts of house prices on fertility decisions, preventing the confounding factor of homeownership, they fail to consider the accompanying movement in homeownership change. Most studies on the demographic impacts of housing price fluctuation focus on a single variable and control for the other, implicitly assuming the choice of housing independent from childbearing.[1] Emphasized earlier, childbearing is not extraneous to housing status and so for the other. Major changes in human life course such as family formation, career building, and childbearing are all likely to raise the possibility of entering homeownership and settling down for the subsequent needs of mental, material, and spatial accommodation. The demand for children can reinforce the demand for more and better housing service and even the desire of moving up the housing ladder in the long term (Clark

[1] This does not mean housing is totally silent from the studies of fertility decision or the other way around. A common argument presumes housing space and children are strong complements, and one of the choices is usually set given in analysis.

and Onaka, 1983; Banks et al., 2004). Needless to mention that private-owned homes are in some cultures deemed as a status good to signal a qualified man for marriage in many cultures. There is also the inverse influence. Stable homeownership not only allows households to allocate more resources to other activities but also provides a stable environment for child raising (Aaronson, 2000; Dietz and Haurin, 2003). This could increase the incentive to have a bigger family size. Mulder (2006a,b) well summarizes this complexity between housing, family formation, and fertility rate conceptually.[2] In particular, the requirement on home location, space, quality, and ownership can be seen as a part of the demand for the quality of family life and of the children's upbringing, to which Kulu and Vikat (2007), Clark (2012), and Clark and Lisowski (2018) partly call for attention.[3] Accordingly, Enström Öst (2012) argues for the growing simultaneity of family housing and fertility decisions overtime, showing that the positive correlation of the two actions has become more prominent in the younger cohort in Sweden as the young generations are facing an increasing insecure occupational and financial environment.[4] In the U.S., Clark and Withers (2009) reports a high ratio of birth events were accompanied by moving events.[5] Clearly, homeownership transition should not be taken away from the analysis of fertility if we want to gain a full picture of the impact of housing prices.

Non-homeowning families are vulnerable to price shocks. The narrowing affordability of homeownership due to the credit constraints and the increasing cost would impinge the decision or the timing of the other important life status transitions such as family formation and workforce participation (Mulder, 2006b; Clark, 2012). This yoke is heavy to young couples particularly: they often do not possess equity or a stable income; at the same time, they are right at the junction of their life course to choose whether or not to have child (Courgeau and Lelièvre, 1992; Sobotka et al., 2011). Without equity, newly-formed families in a booming housing market face more difficulty to pay for a mortgage down payment had

[2]We limit the scope of this paper from the impact of transformation in social norms in recent years, which could turn down the demand of housing prompted by other family status transition. About this issue, see the discussions in Mulder (2006a) and Drew (2015). Additionally, the influence of fertility rate on housing prices is assumed away because such channel would hardly be captured in micro-level analysis.

[3]In a very different setting, Sato and Yamamoto (2005) discusses an equilibrium of fertility rate, population density, and urbanization in Japan, which alludes to the importance of the relationship between residence and childbearing.

[4]The caveat of this positive relation is that homeownership reflects the fulfillment of the desired housing space. If homeownership per se instead of its incurred housing service becomes the goal, the relationship may not hold.

[5]The effect of the status of having a child and the event of childbirth are different. It is the event that would raise the likelihood of moving. The status of having children decreases the likelihood empirically perhaps because children fortifies family and reinforces the desire of sedentariness (Clark, 2013).

their income remains unchanged, which became even more notable in the era of a tighter housing mortgage market after the Financial Crisis (Lennartz et al., 2016). If the social norm strongly regards homeownership as a requirement of formal family formation, the rising financial burden from the growing housing cost would drain their savings and further crowd out the resources that could be used elsewhere, such as parenting, as observed in some parts in Europe and East Asia.[6] High homeownership rates under this circumstance would not reflect that the pervasive positive wealth effect caused by a housing price increase but instead stresses the heavy burden of housing. In Southern Europe, where such preference for homeownership is strong, insufficient housing rental market and inferior access to housing mortgage accompanying with a high homeownership rate are likely attributable to the extreme-low fertility rates (Mulder and Billari, 2010). Lo (2012) also observed a negative relationship between homeownership and fertility in a cross-county study on Taiwan. Indirectly yet untested, Turner and Seo (2007) suggests this possible substitution observed from the U.S. data.

In a microeconomic analysis, if we regard children as an economic good, the decision on home buying and childbearing should not be considered in a static framework as but of them are durable.[7] Families desiring a homeownership and a child contemplate not only their current but future satisfaction from the housing service and children's development, and the expectation of future prices determines their willingness to take action today. Both the current price level and the expected price are in family's consideration. A large literature about housing and fertility choice have discussed about their respective dynamics (e.g. Hotz et al. (1997); Ortalo-Magne and Rady (2006)). In addition, buying a home is not an action to which perfect capital market can be unconditionally assumed (Daysal et al., 2020). A potential buyer must accumulate enough savings in practice to pay at least part of the value of home.[8] To achieve that, families may alter the preferred time of childbearing to achieve their financial goal. Such lagged effect of home prices is the same type of the regular

[6]Though the homeownership rate is falling among the younger cohorts in the U.S., it is not a decisive evidence for the preference change of American young families. See Myers et al. (2020) for the discussion.

[7]The Beckerian setting is to assume the utility of parent comes from not only the number of children but their "quality." The discussion of the effect of home price on fertility in this paper intentionally circumvent the question of the demand of children's quality to prevent digression. (For the discussion, see Becker and Lewis (1973).) Different from the effect of average permanent income, the variation of home prices does not involve in a potential change on the opportunity cost of mother's time nor the expected return of the child's education investment (Becker, 1960, 1965). Moreover, this inverse effect on fertility through quality demand change may not yet be dominating in individual-level analysis on American data, as Black et al. (2013) discovered a positive relationship of husband income and fertility using the 1990 U.S. census data.

[8]In the U.S., a conforming home mortgage with down payment ratio smaller than 20% requires the borrower to buy a mortgage insurance.

tempo effect caused by rising female education and average income (Lutz and Skirbekk, 2005), except it is purely due to the economic reason. A high house price implies a tougher barrier of the credit constraint that bars families from their unconstrainted optimal timing of homeownership transition and childbearing. The discretion in family finance and the possibility of intertemporal choice on home buying and childbearing reflect the importance of a dynamic framework for the co-movement to understand the interplay of the two family choices.

3 Research Strategy

3.1 Conceptual Framework

In a nutshell, homeownership and childbearing are not mutually independent decisions for family, and house price has an impact on both simultaneously. Due to the credit constraint and the expectation of future prices, the optimal timing of the choices could be determined by the current price level and its trend. To examine this co-movement, we extend the common single-causality framework to a dynamic two-dimensional choice question: a non-homeowning family considers home buying and childbearing together in each given time period. Without equity holding nor extraordinary wealth, the family needs to face the credit constraint on home purchase and adjust its optimal decision based on the prices available. The current price level determines the current affordability, and the recent price variation affect the expectation of future house prices. An increase in local house price would render house buying more difficult and shape family's belief. A short-term variation may be a transitory deviation. A middle-term change may reflect the trend.

The family does not take the decision once for all in its life course. Before becoming homeowner or realized desired family size, it keeps updating the relevant factors and act optimally following the current condition. In each time point, the choice is made between the crossing of the two yes or no questions: to buy a home now? To have a baby now? If the desired family status is yet to be satisfied, the question will be reconsidered later once the information is updated. In the exercise we track the behavior of the non-homeowners, so the statistics would reflect the state transitions and their association with home price changes.

The extension does not mean a great jump to the comprehensive knowledge of the family behavior. To keep it simple, this framework limits the choice set in each dimension to be a simple binary one, and the quality and quantity of housing services and children are precluded. It is just an abstract of a general decision between homeownership and parenthood.

Many other determinants are also ignored. Unobservable and excluded endogenous factors, such as credit score, abortion access, and migration, all take a part in the determination of home buying and childbearing. We acknowledge the extent of interpretation power of this framework is limited. In any case, as far as one does not overgeneralize the interpretation of the model, it provides some insights on family behavior regarding homeownership and childbearing across time for the investigated group.

3.2 Empirical Approach

The main purpose of the empirical exercise is to explore the proposed effects of house price level and changes on the joint behavior of childbearing and entering homeownership and childbearing, deemed as joint dynamics in the family life course at an individual level. To this end, we examine the relationships between the observed local house price levels and variations and the probabilities of women giving a birth and entering homeownership in a single model. Given the distinctive nature of the two family behaviors, as will be discussed soon, and the assumption on adaptive expectation of family budget, we apply a multinomial logit (MNL) model for the estimations.[9]

We model homeownership and parenthood of a new child as two binary choices. The crossing of the two binary choices creates four choice alternatives with no natural ordering at each time point. In line with the conceptual framework, we consider that a non-homeowning woman f, representing her family, decides what to do among the four alternatives in each time period t. In the next time period, the information of the local housing market is updated, and she makes a new choice. Each choice is considered a decision-making case n such that $n \in \{f \times t\}$. In each case, the utility the woman would obtain from the alternatives follows the standard random utility model (RUM), and she makes her choice by choosing the one that would give her the greatest utility. The RUM formulates the utility of case n from alternative j as $U_{nj} = V_{nj} + \varepsilon_{nj}$, where V_{nj} is called the representative utility which is a function of the observable factors, and ε_{nj}, the disturbance component, captures the utility that is influenced by the other unobservable factors. Because the alternative i will be chosen if and only if $U_{nj} > V_{nj}$ or $\varepsilon_{nj} < \varepsilon_{ni} + V_{ni} - V_{nj}$ for all $j \neq i$, the probability of the choice is the product of all the cumulative distribution of ε_{nj} under $\varepsilon_{ni} + V_{ni} - V_{nj}$ for all $j \neq i$ given ε_{ni}. We assume the all the disturbances are independently and identically distributed ($i.i.d.$)

[9] The basic settings of the logit model for multinomial discrete choices are well documented in the econometric literature, including Ben-Akiva and Lerman (1985),Imbens and Wooldridge (2007), Train (2009), and Walker and Ben-Akiva (2011). This paper mostly follows the terms that are used in Imbens and Wooldridge (2007) and Greene (2012).

and following Gumbel type-1 distribution, and we can derive the closed-form expression of the probability of the choice in the following equation, which stands for the logit choice probability.

$$\begin{aligned} P_{ni} &= Pr(\varepsilon_{nj} < \varepsilon_{ni} + V_{ni} - V_{nj} \ \forall j \neq i) \\ &= \int \left(\prod_{j \neq i} F(\varepsilon_{ni} + V_{ni} - V_{nj}) \right) f(\varepsilon_{ni}) d\varepsilon_{ni} \\ &= \int \left(\prod_{j \neq i} e^{-e^{-(\varepsilon_{ni}+V_{ni}-V_{nj})}} \right) e^{-e^{-\varepsilon_{ni}}} d\varepsilon_{ni} \\ &= \frac{e^{V_{ni}}}{\sum_j e^{V_{nj}}}. \end{aligned} \tag{1}$$

For each case, we observe a vector of factors, x_n. It consists of the woman's information at the time, including her demographic characteristics and financial conditions, the local economic indices, and the local house price index. Because the two binary choices are of different nature (homeownership and parenthood do not have accurate measures in common save the pecuniary costs), there is no common variable that varies among the alternatives.[10] Thus, the model assumes no alternative-specific variable is considered in the model and all the observable variables are individual- or case-specific.

For the purpose of empirical implementation, we assume V_{nj} as a linear function of x_n with a vector of choice-specific parameters, β_j. The convenient specification is defined as

$$V_{nj} = \equiv x_n' \beta_j = \beta_{j0} + \beta_{j1} p_{st} + z_n' b_{1j} + \varphi_{st} b_{2j} + v_{sj} + w_{tj}, \tag{2}$$

where in our benchmark model p_{st} denotes the real-term median local market house price level in MSA s in year t, z_n denotes the vector of case-specific variables of case n for woman f in year t, and $varphi_{st}$ denotes the vector of local economic factors collected. The vector z_n includes the woman's age, race, education level, partnership (marriage or cohabitation) status and change, employment status, and the current total family income. The vector $varphi_{st}$ includes the state-level unemployment rate and MSA-level personal income per capita and an index for nationwide recession. The potential locational fixed effects and the year effects are captured by variable v and w, which represent the census division invariant

[10] As the theoretical framework links the two behavior through the family budget, the consumption of a composite of all other goods, or inversely the cost of the actions, is a legitimate alternative-specific variable. However, we cannot observe the price of homeownership a woman faced nor the accurate cost of childbearing from the data.

effect and the five-year group invariant effect.[11] Moreover, as it is almost certain that the error terms are not independent for the same individual throughout the panel, we implement clustering of the standard errors at the individual-level throughout the estimations. Since we are also interested in the effect of house price variation, we replace the price level with Δp_{st} and $\%\Delta p_{st}$ in the alternative reduced-form model, which represents the two- or four-year price change and growth rate in real term in MSA s of year t.

Regrading with the parameters, the constant term, β_{j0}, and the parameter of our main interest, β_{j1}, are scalar, and b_{1j} and b_{2j} are two sub-vectors in β_j. The alternative $j = 0$, denoting the choice that the woman would not enter homeownership nor have a child, is set to be the base outcome and β_0 is normalized to a vector of 0. The interpretation of the estimates is however not very straightforward. By Equation (1), the value of a parameter β_{ik} directly expresses the marginal effect of the k-th variable x_k on the natural logarithm of the relative probability for alternative i, RP_{i0}, which is defined as the proportion of the probability of i to the probability of the base outcome ("doing nothing"). The following equation states this relation for case n.

$$\ln RP_{ni0} \equiv \ln\left(\frac{P_{ni}}{P_{n0}}\right) = x_n'\beta_i \tag{3}$$

$$\beta_{ik} = \frac{\partial \ln RP_{ni0}}{\partial x_{nk}}. \tag{4}$$

Though a positive β_{ik} indicates a positive marginal effect on the log-relative probability, it does not necessarily imply a positive marginal effect on the probability (Greene, 2012). In a MNL model, the marginal effect of x_k on the probability of choosing alternative i is a function of the probability of choosing i and all the estimated parameters. Equation (5) shows the calculation. Later, we mainly report the estimated marginal effects on the probability to give a more intuitive interpretation of our findings.

[11] This setting of the fixed effect groups, which differs from the common state or city fixed effect and year fixed effect, is designed due to the lower limit of choice sample size for estimation.

$$\begin{aligned}
\frac{\partial P_{ni}}{\partial x_{nk}} &= \frac{\partial \left(e^{x_n'\beta_i} / \sum_j^J e^{x_n'\beta_j} \right)}{\partial x_{nk}} \\
&= \frac{e^{x_n'\beta_i}}{\sum_j e^{x_n'\beta_j}} \frac{\partial \left(x_n'\beta_i \right)}{\partial x_{nk}} - \frac{e^{x_n'\beta_i}}{\left(\sum_j e^{x_n'\beta_j} \right)^2} \left(\sum_j e^{x_n'\beta_j} \frac{\partial \left(x_n'\beta_j \right)}{\partial x_{nk}} \right). \\
&= P_{ni} \left(\beta_{ik} - \sum_j P_{nj}\beta_{jk} \right) \equiv P_{ni} \left(\beta_{ik} - \bar{\beta}_{ik} \right)
\end{aligned} \tag{5}$$

The MNL model is estimated by using the maximum likelihood method. Back to (1), this model naturally requires the outcomes to be exclusive, exhaustive, and finite and satisfies the property of independence from irrelevant alternatives (IIA). It is a consequence of the assumption of *i.i.d.* disturbance. We argue it is reasonable to assume the outcomes more or less meet the requirements.[12] As they are the crossing of two binary choices, the first three properties are automatically satisfied. The satisfaction of IIA property is more debatable since, had we limited choice from buying home alone or giving birth alone, the predicted odds of the rest alternatives may not remain the same, especially when with a big set of variables is considered. Nevertheless, we show that the estimation of the main parameters of interest pass the Hausman tests for IIA property for all combinations of alternatives with the base outcome. Also, applying to a panel data, we implicitly assume the disturbances, as well as the unobserved factors, are independent over time. It is, again, a simplistic assumption in compromise for the convenience of the model estimations.[13] One related underlying assumption is that women would adapt their optimal path in each time period given the new state variables. The idea *per se* is very similar to discrete-choice dynamic programming except that it contains irreversible state transitions and the only state variable connecting period is the amount of private asset (Keane et al., 2011).

A few more things are taken into consideration in this model. First, Mulder (2006b) states the possibility of inter-correlation between housing prices and local fertility rate. Though an individual decision can hardly affect the whole MSA's housing price level, it is nonetheless a threat to identification. Local housing prices correlate with local macroeconomic conditions. The state-level unemployment rate and personal real income are for this sake introduced

[12] As for the choice of a logit model, according to Amemiya (1981), the estimations under the logit assumption does not visibly differ from one's under the assumption of normal distribution (a probit model). Small et al. (2007) argues that the advantage of the MNL in its simplicity outweighs the cost of the assumptions.

[13] For example, family wealth and health condition of the family member are likely to be autocorrelated but hard to be observed due to the data limitation.

to control for the macroeconomic variation of the region. Second, we limit our samples in the women who did not move across MSAs during the time of the tracked house price change (two or four years). This is to prevent the endogeneity problem of the movers who choose the place where the housing market is preferred. Last, the estimated standard errors are calculated using the sandwich estimator in order to be robust against the unspecified heteroskedasticity.

4 Data

We construct an individual-level panel dataset of women in non-homeowning and independent families with the local house price and other economic index in order to investigate the effect of the house price variations on the family behavior.[14] Our main data sources are the restricted-used Panel Survey of Income Dynamics (PSID) and the Cross-National Equivalent File (CNEF). The local house price data is built from the MSA-level Housing Price Indices (HPI) from the Federal Housing Financial Agency (FHFA) and the Longitudinal Tract Database (LTDB). Other supplementary data for the local and national economic performances comes from multiple resources including Bureau of Economic Analysis (BEA), the Local Area Unemployment Statistics (LAUS), and Federal Reserve Economic Data (FRED). All the monetary measures are in real term, inflated to 2011 dollars using the CPI for All Urban Consumers (CPI-U).

The PSID is a public longitudinal survey on the financial conditions of U.S. families conducted by the University of Michigan since 1968. It drew a group of families in the first survey and then follows those families and their descendants and records their financial and demographic information including moving, homeownership, and childbearing every one or two years.[15] Its restricted-used version provides the geographic information of the observations. This allows us to pin down the respondent's residence and link to the local economy. This advantage makes tracking family status transition and its relation to the local housing price level possible. We take the sample from the surveys from 1985 to 2015,

[14]Though a great proportion of our sample are presumably first-time home buyers, we refrain to use the term because some women in the sample are reportedly living in an owner-occupying unit initially and then moved out.

[15]The PSID was initially an annual survey with detailed financial variables (especially regarding the family wealth) collected every four years. Since 1997, the PSID survey became biennial. A group of Latino and immigrant families were later added into the survey. In our analysis we exclude the Latino families added in 1990 and 1992 because the PSID does not assign proper weights to them.

in total 22 waves.[16] Our sampling strategy imitates the work of Lovenheim and Mumford (2013) at a certain degree. Women aged 20 to 44 in the financially independent family who are either the family head or the partner (spouse or cohabitator) or the head are selected. This choice is based on the common childbearing period (age below 45) and the likelihood that the respondent (or her partner) is financially independent. There are in total 77,792 such observed cases in PSID. As we are interested in the behaviors of family which are facing the dual decisions, we further limit our sample to be the group who live in the area where the local HPI are available, are not a homeowner, and did not move to other MSAs during the time window of house price change considered (2 or 4 years).

The original PSID data structure is mostly family based. To construct an individual level panel, we borrow the data framework of the CNEF-PSID. The CNEF is a research project organized by the Ohio State University, aiming to construct a uniform international social and economic data sets. Its PSID branch publishes a processed individual-level panel data set of the PSID up to year 2015 with a limited number of variables. Although it only contains limited information, it serves as the backbone of a comprehensive individual-level data for our purpose. With the help of the WZB-PSID tools developed by Ulrich Kohler, we merge the PSID data with the CNEF-PSID data set and build our main data set.[17] In the analysis, we apply the standard cross-sectional PSID weight constructed by CNEF-PSID. The standard weight provided by the PSID accounts for the original family's national representativeness and attrition over time.[18] In estimation, the case weight is the individual's PSID weight divided by the number of cases of the individual in the sample to prevent the over-representation issue. One feature of the PSID weight is that it excludes women who appeared in the sample by marrying in or cohabitating with the core PSID members. This setting avoids data attrition due to divorce or cohabitation break up, but also causes the loss of a considerable number of observations. In the section of alternative specification, we construct a supplementary weight (the "extended weight") to include these women in estimation by assigning them the same weight from their partner. As will be discussed later, adding these sample would affect the results only mildly.

[16] Some key variables were added into the survey only since the wave of 1985, and the CNEF has not cover the 2017 data by the date of writing.

[17] We thank to David Brady and Ulrich Kohler for their help to reconstruct the PSID data for the analysis with the WZB-PSID tools and relevant commands they developed.

[18] The broad idea of the PSID weight is that the members and the descendants of the original surveyed families (the "core members") are assigned a weight that reflects the possibility of the family being selected in 1968. Then, accounting for the conditional probability of attrition, the weights of remaining respondents grow slightly in every survey.

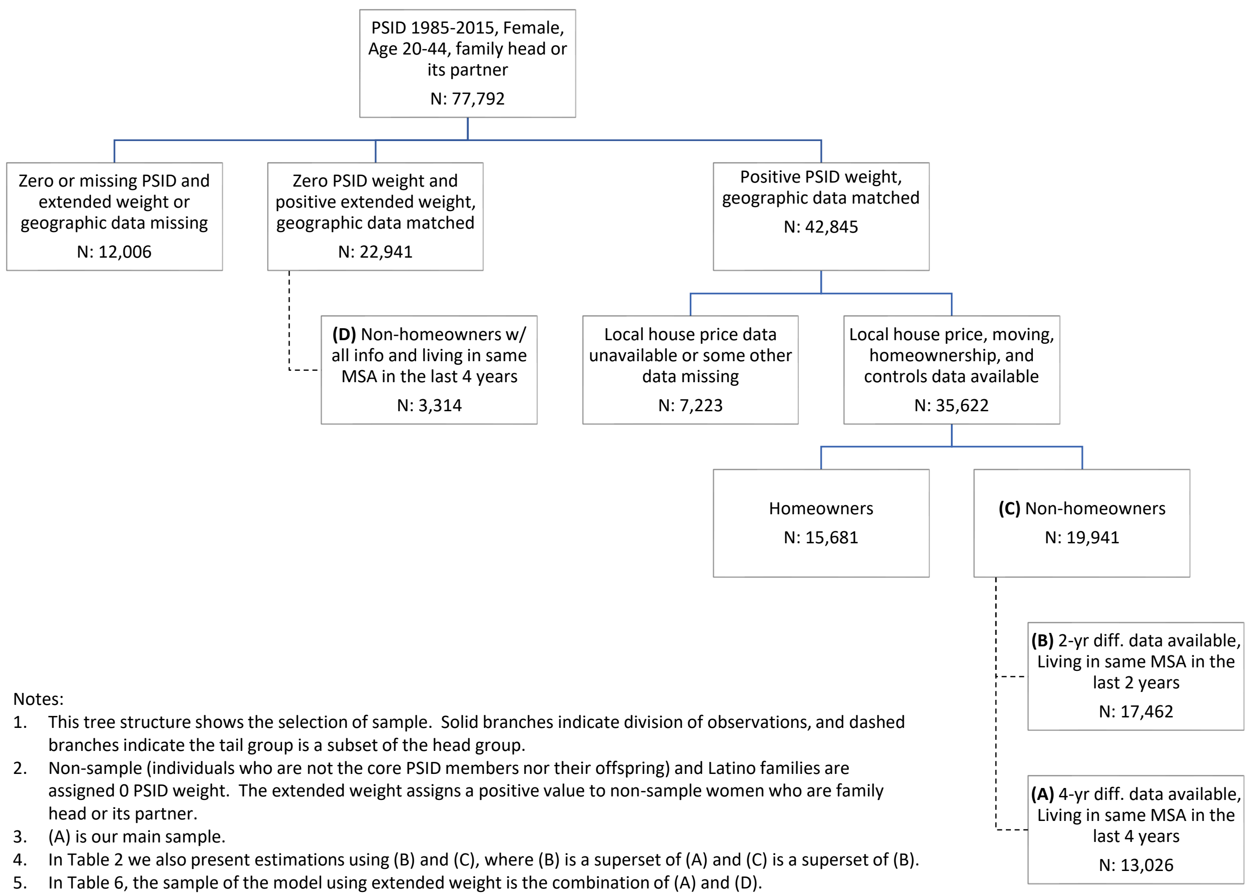

Figure 1: The tree structure of PSID sample selection.

After censoring PSID weight and (cross-sectional) data completeness, 35,622 cases are left. Among them, 19,941 non-homeowner cases are available for analysis. To concentrate on the effect of housing and childbearing, only women who did not move inter-MSA in the past four years are kept, which ends up with the size of 13,026. This is our main sample. Figure 1 shows the method of sample selection, and the main sample is presented as box (A). In this sample, the unconditional probability of entering homeownership only is 6.9%, of childbearing only 6.7%, and of doing both in the same time period 0.6%. Later, we also estimate models using sample box (B) and (C) to examine the representativeness of our main sample. Moreover, box (D) is added to our main sample when we use the extended weight.

Table 1 presents the summary statistics of the sample with positive PSID weight and no missing information. Except the first three rows (the dependent variable) and the last three rows (the independent variable of interest), all the other variables are introduced into the model as controls. Column (1) summarizes the whole available sample, regardless whether the women are existing homeowner or have moved from other MSA recently. This sample is represented by the right most box of the third tile in Figure 1. Column (2) restricts the sample to only non-homeowners, which is indicated as box (C) in the figure. Column (3) summarizes the main sample, namely the box (A) in the figure. There is a clear demographic difference between the groups.

Comparing column (1) and (2), the non-homeowners are in general younger, with lower education level, much less likely having a partner, and with a higher rate being black, while the unconditional likelihood of giving a birth is almost the same with the whole sample. Not surprising, they also tend to have a lower family income, in average \$42,572 versus \$65,323 annually.[19] They are more vulnerable to the house price growth not only because of the lower income but also because they do not possess any equity hedge. The demographic difference between column (2) and (3) is much smaller, except the average age of non-migrant women is similar with the whole sample. Notably, women who stayed in the same metropolitan area in the last four years have a lower probability of entering homeownership and birth, probability because their family income is in average lower and they are more likely to like alone. This suggests a relation between migration and housing and fertility. Though this is not what the focus of this paper, it is a fact that deserves more attention.

Our outcome variable for the regression model is defined as four mutually exclusive alternatives of actions in a time window. It is set as the time period between the last pair surveys, and we track the survey dates to month. During each time window, a female

[19]The distribution of total family income is right skewed. The medians are both lower.

Table 1: Summary statistics of the PSID sample.

Category	Variable	(1) All available cases with full info		(2) Non-homeowners		(3) Non-homeowners and non-migrants	
		Mean	S.D.	Mean	S.D.	Mean	S.D.
Decision	New homeownership only	0.079	(0.270)	0.076	(0.265)	0.068	(0.252)
(in Prob.)	New birth only	0.074	(0.261)	0.074	(0.262)	0.067	(0.250)
	New ownership and new birth	0.009	(0.092)	0.008	(0.087)	0.006	(0.075)
Demographic	Age	33.766	(5.983)	32.125	(5.945)	33.603	(5.470)
	Partnership	0.586	(0.493)	0.391	(0.488)	0.359	(0.480)
	White	0.528	(0.499)	0.397	(0.489)	0.330	(0.470)
	Black	0.415	(0.493)	0.545	(0.498)	0.618	(0.486)
	Other race	0.057	(0.232)	0.058	(0.233)	0.051	(0.220)
Education	High school diploma	0.442	(0.497)	0.489	(0.500)	0.533	(0.499)
	Some college	0.289	(0.453)	0.309	(0.462)	0.307	(0.461)
	College graduate	0.270	(0.444)	0.202	(0.401)	0.160	(0.367)
Family size	No Child	0.293	(0.455)	0.342	(0.474)	0.277	(0.447)
	One Child	0.243	(0.429)	0.245	(0.430)	0.246	(0.431)
	Two Children	0.277	(0.448)	0.228	(0.420)	0.255	(0.436)
	More than Two	0.187	(0.390)	0.184	(0.388)	0.222	(0.416)
Fianacial	Employment Status	0.812	(0.391)	0.798	(0.401)	0.783	(0.412)
	Real family income in $1,000	65.323	(64.829)	42.572	(41.348)	40.730	(41.675)
Housing	2-Year Price change in $1,000					2.941	(29.836)
Market	4-Year Price change in $1,000					5.073	(49.219)
	Median price in $1,000	176.984	(89.604)	181.891	(95.517)	181.085	(96.106)
Observations		35622		19941		13026	

Note: The sample are women who are either family head or its partner, age between 20 and 44, with positive weight and all information available from PSID 1985-2015. Non-homeowners is equivalent to the Sample box (C) in Table 1, and non-homeowners and non-migrants are equivalent to Sample box (A) in Table 1, where non-migrants means people who stayed in the same MSA in the last four years. All monetary means are inflated using the CPI-U in real 2011 dollars.

respondent chooses either to do nothing, buy a home, give a birth, or both. There is however a doubt for the setting. Since the PSID shifted from annual to biennial in 1997, the time window for respondents after year 1997 became one year longer than whom were surveyed before. To examine whether this would become a confounding factor, we set another group of estimations that fix the time window of the outcome variable to the period between current survey and the survey taken two years ago, and regress with the respondents from surveys of the odd years only. The outcomes are presented in section 5.3. We show that the results from two groups are qualitatively same, with the estimations from the second group inevitably suffer from higher standard errors. For convenience, we call the outcome variable in the standard group the "flexible window output" and in the alternative one the "fixed window output" in the following sections.

The house price data of the MSAs are imputed from two sources. On the one hand, the FHFA publishes the quarterly Housing Price Indices (HPI) of 403 MSAs.[20] The earliest recording dates in 1976, though most of the series start from 1980s. These indices estimate the longitudinal trend of the price level of local single-family houses using both repeat-sales prices and appraisal data (not seasonally adjusted). Except eight MSAs, all indices set 1995 as the base year. On the other hand, the Longitudinal Tract Database (LTDB) from the Brown University has the data of the cross-sectional tract-level median home values, which it calculated from the decennial Census. Since the HPIs do not represent cross-sectional price differences between cities, we take the cross-sectional home value data of year 2000 from the LTDB and calculate the average median home value of each MSA, then together with the FHFA HPI we construct a panel of imputed yearly average house prices. Although the LTDB also has the median rent data, it is unfortunate we cannot find a reliable local longitudinal information for rents.[21]

The last three columns in Table 1 shows a summary of the house price changes in real term.[22] In average, the house prices experienced a net growth in the past 30 years, despite a huge slide between 2007 and 2011. For our main sample, the net average two-year real house price change is \$2,941 and of four-year change is \$5,073. The high standard errors partly reflect the fluctuations in time series and partly reflect the huge diversity of house

[20] The list of the MSAs changes over time due to the demographic change. We update all the locational information in accordance to the September, 2018 Delineation of the United States Office of Management and Budget (OMB).

[21] The FHFA recommends the CPI-U of all items less shelter for estimating the inflation of HPI. However, a pilot estimation shows it does not produce a notable change on the results.

[22] The statistics of house price changes for column (1) and (2) are suppressed because, without dropping sample who migrated, migration made house price change endogenous.

price growth between cities. Even at the census division level, this spatial difference can be easily spotted out by comparing the distributions of house price change. As Figure 2 indicates, the local house variations in the coast areas are much greater than the cities in the Mid-west and the South for households in PSID.

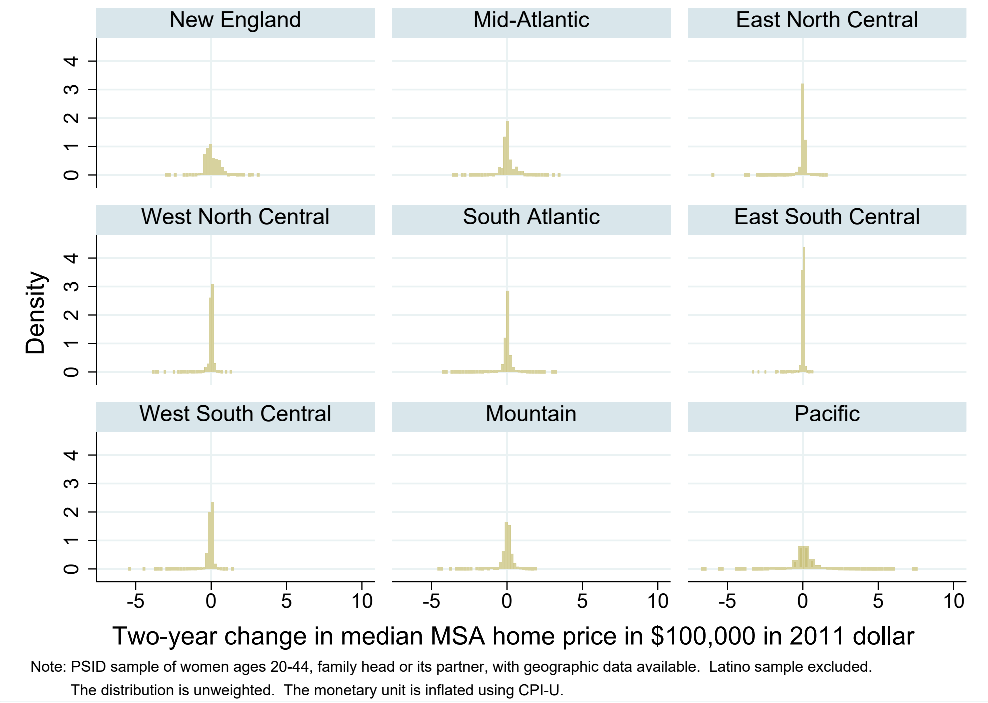

Figure 2: The distribution of 2-year home price change between 1985 and 2015, by census division.

Other supplementary data sources provide aggregate level data to control for the regional or national macroeconomic conditions. We take the MSA-level average personal income from the BEA, the state-level unemployment rate from the LAUS, and create a yearly national recession index by taking the annual average of the quarterly recession index from the FRED. These variables reflect the state economic performance and the broad national economy health about which families are likely to be concerned when they formulate the expectation for the future market condition.

5 Results

Our benchmark statistical model presented below takes the flexible window output and the PSID standard weight. Without further specification, the sample is women who did not move inter-MSA in the past four years, namely the sample box (A) in Figure 1. Most tables in this section reports the estimated marginal effect to give an intuitive and comparable interpretation. It should be reminded that the predicted effects are quantitatively meaningful only at the margin of the change.

5.1 The marginal effect of variables of interest

Table 2 reports the main results of the estimated average marginal effects of the house price value and variations from the MNL model earlier described.[23] Each column presents the estimated marginal effects and their standard error of the specific independent variable on the three alternatives, with no action as the base alternative with all coefficient normalized to 0. The upper panel reports the results from the benchmark model. The sample is limited to women whose residential data in the past four years is available and records no change in residential MSA during that time, namely the sample (A) in Figure 1. All estimations include the full set of controls listed in Table 1, regional fixed effects at the census division level, and time fixed effects at 5-year level. The standard errors are estimated by the sandwich estimator clustered at the individual level. Column (1) shows the marginal effects of the real house price level, column (2) and (4) show the effect of the two-year and four-year real price change, and column (3) and (5) show the effect of the two-year and four-year real price growth rate. The bottom panel reports the estimation results using more general sample selection rules for comparison. The model for the left three columns slackens the rule of selection to two-year data availability and no residential MSA change. This adds the sample size by more than four thousand. The model for the rightmost column includes all women who regardless whether she moved from another MSA in the past, further adding two thousand observations. They are represented by box (B) and (C) in Figure 1.

The estimation on the effect of the real local house price level shows that families living in an expensive area generally has a lower chance to enter homeownership but a higher chance to have a new child at the renting stage. A \$100,000 difference in the local house price unsurprisingly leads to 4.6 to 5.2 percentage points decrease of the probability of entering homeownership, which is in line with earlier work using PSID (Henretta, 1987). Oppositely, it contributes to 1.0 to 1.7 percentage points of the probability of women giving a birth. Regarding to the doting both alternative, it seems that the discouragement to

[23]The MNL model passes the IIA tests for all the parameters of the main variables of interest. Table A1 shows the test results. Moreover, because the baseline probability of the alternatives varies in a wide range, elasticity does not provide intuitive interpretation. Because of this reason, we report the marginal effects only. Table A2 reports the elasticity table for the models in the upper panel of Table 2.

Table 2: The MNL model estimates of marginal effects of house price level and variations.

Dependent Choice	(1)	(2)	(3)	(4)	(5)
	Independent Variable				
No inter-MSA move in the past four years					
	Median House Price ($100,000)	2-Year Price Change ($100,000)	2-Year Price Growth Rate	4-Year Price Change ($100,000)	4-Year Price Growth Rate
Homeownership only	-0.046***	-0.031	-0.137*	-0.016	-0.062
	(0.0143)	(0.0254)	(0.0747)	(0.0159)	(0.0426)
Birth only	0.017***	-0.007	0.011	0.004	0.021
	(0.0059)	(0.0126)	(0.0379)	(0.0085)	(0.0231)
Both	-0.011***	-0.007	-0.029	-0.013***	-0.034**
	(0.0039)	(0.0058)	(0.0211)	(0.0045)	(0.0160)
N	13026	13026	13026	13026	13026
No inter-MSA move in the past two years				Regardless moving	
	Median House Price ($100,000)	2-Year Price Change ($100,000)	2-Year Price Growth Rate		Median House Price ($100,000)
Homeownership only	-0.052***	-0.010	-0.087		-0.047***
	(0.0136)	(0.0191)	(0.0574)		(0.0127)
Birth only	0.012**	-0.015	-0.023		0.010**
	(0.0049)	(0.0090)	(0.0272)		(0.0043)
Both	-0.007	-0.007	-0.032		-0.007
	(0.0050)	(0.0081)	(0.0235)		(0.0044)
N	17462	17462	17462		19941

Note: Models differ only in the independent variable of interest. All estimates includes controls for partnership, race, number of children, total family income, and employment status, age group dummies, educational attainment dummies, state-by-year unemployment rate, MSA-by-year real income per capita, national recession index, and geographic (census division) and time (five-year period) fixed effects. Robust standard errors in parentheses account for clustering at individual level. Region (census division) and time (five-year) fixed effects in all specifications. Significant at *** p <10%, ** p <5%, * p <1%.
Source: Non-homeowning women who are either family head or its partner, age between 20 and 44, with positive weight and all information available from PSID 1985-2015.

homeownership slightly outweighs the encouragement to childbearing, leading a 0.5 to 1 percentage decrease in probability. Yet the standard errors are relatively high as a result of the small size of observations on this choice. The results of the first two alternatives are robust as dropping the last alternative does not make a substantial drift of the estimated values (results not shown here). To give a sense to the number, the difference between the median sales prices of houses between 2011Q1 and 2015Q1 is $50,163 in 2011 USD, according to the Federal Reserve. If we take the estimates from the upper panel for granted, this price level difference by itself would generate a 2.3 percentage points decline, or about a 33.3% decrease, in the likelihood of homeownership transition only; a 0.9 percentage points increase, or 13.4% growth, in the likelihood of childbearing only; a 0.5 percentage point decline, or 83.3% decrease, in the likelihood of doing both, if other things remain equal.

On the other hand, we only find weak evidence in support of the marginal effects of house price change. The results posted in Table 2 suggest that experience of house price growth can deter women from entering homeownership, as the estimates of the effect on both entering homeownership only and doing both are negative. Consistent estimates notwithstanding, the suggestive effect is only statistically meaningful for four-year price change on doing both. The estimates from column (4) indicates a $100,000 price increase would result to 1.3 percentage points decrease in the probability of the choice, and the estimates from column (5) alludes a 40% increase in house price would lead to an effect at the same magnitude had the effect been linear to price growth rate, *ceteris paribus*. The estimates of marginal effects on childbearing are weak and mixed. For two-year price change and growth, the estimates are small and inconsistent in direction with large standard errors. For four-year price change and growth, the estimates are consistently positive but very weak and lack of statistical power. However, putting together the with the estimated effect of doing both, these results signify a net decline in childbearing likelihood.

These results imply an interesting interaction between house price variation and the dynamics of homeownership and childbearing, and at the same time goes along with the literature. To start with, Lovenheim and Mumford in their 2013 paper shows that argue the increasing trend in fertility in early 2000s is likely contributed by the wealth effect of homeowners due to the home equity appreciation, while neither house price level nor change surge are statistically associated with the childbearing likelihood of renters.[24] Our estimations indicate instead that high price leads to a tradeoff between homeownership and

[24]The focus of Lovenheim and Mumford (2013) is the fertility behavior of homeowners. In their analysis, they present the regression results of renters as the side finding to compare with their main results. Here we compare our results with their side finding.

childbearing for those families, or at least they are compelled to postpone homeownership and switch the orders of homeowning and parenting in their family life course. To link this to Lovenheim and Mumford's results, we estimate the effects of house price level and change on childbearing likelihood only using a simple linear regression model with the same set of controls plus year and state fixed effects and report the results in the upper penal of Table 3. In the first three columns, the model used regresses the probability of woman giving a birth between the last two surveys, regardless whether she entered homeownership during the same time period. State fixed effects and year fixed effects are controlled. This is in line with the spirit of Lovenheim and Mumford's main model. Though detailed settings and variable definitions are different, we obtain the result rejecting the relationship between house price and net childbearing likelihood, which are in congruence with their finding on renters.

The model for column (4) to (6) in the upper panel regresses the probability of woman giving a birth only between the last two surveys, excluding incidents of entering homeownership and giving a birth during the same period, and in the bottom panel the model regresses the other two alternatives on the same set of independent variables. These regressions show that local house price has a prominent positive relationship with the probability of childbearing only and a negative relationship with the probability of entering homeownership as well as doing both in the same time period. Putting them together, the evidence suggests a consequential delay of families entering homeownership caused by high local house prices, whereas the course of parenting is likely to take place at any rate, and even with a minor increase in the net probability. A $100,000 higher in local house price relates to a 1.8 percentage points increase in the probability of childbearing only and a 1.3 percentage points decrease in doing both in the same period, resulting a 0.5 percentage point net increase in the probability of childbearing, which reflects a 7.5% growth.

Financial constraint is the most plausible mechanism to explain this outcome. Because of the larger amount of required mortgage down payment and a higher expected mortgage payment following, prospective home buyers are apt to prefer to wait longer before entering the market. Meanwhile, childbearing decision seems merely affected by the price shock, and, if any, is probably due to the temporarily loosen family budget since the expense on the mortgage is postponed. This suggests more children will be born into non-homeowning families, though it does not necessarily imply a smaller housing space for children as these families can still expend their housing space by moving to a bigger rental unit. The underlying limitation of the estimates here is that it only covers women who did not move across

Table 3: The linear probability model estimates of marginal effects of house price level and variations.

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Choice	Independent Variable					
	Median House Price ($100,000)	2-Year Price Change ($100,000)	4-Year Price Change ($100,000)	Median House Price ($100,000)	2-Year Price Change ($100,000)	4-Year Price Change ($100,000)
Birth	0.005 (0.0079)	0.002 (0.0142)	0.001 (0.0087)			
Brith only				0.018*** (0.0066)	-0.003 (0.0125)	0.007 (0.0077)
N	13026	13026	13026	13026	13026	13026
Homeownership only	-0.036** (0.0143)	-0.041* (0.0241)	-0.016 (0.0150)			
Do both				-0.013*** (0.0045)	0.005 (0.075)	-0.006 (0.0042)
N	13026	13026	13026	13026	13026	13026

Note: Models differ only in the independent variable of interest. All estimates includes controls for partnership, race, number of children, total family income, and employment status, age group dummies, educational attainment dummies, state-by-year unemployment rate, MSA-by-year real income per capita, national recession index, and geographic (census division) and time (five-year period) fixed effects. Robust standard errors in parentheses account for clustering at individual level. Region (census division) and time (five-year) fixed effects in all specifications. Significant at *** p <10%, ** p <5%, * p <1%.
Source: Non-homeowning women who are either family head or its partner, age between 20 and 44, stayed in the same MSA in the last four years, with positive weight and all information available from PSID 1985-2015.

MSAs through the survey period.

The results in Table 2 and 3 both suggest no strong relationship in general between short-term house price change or growth and the two family decisions, hinting that such short shock itself is not really an influential factor. One possible interpretation is that short-term price change may not be enough to let families give up or delay homeownership, as house price variations can be transitory. If the growth persists longer, an expectation of continuing growth trend may be formed in public. It starts to have an impact on the costliest choice, doing both at the same time. Thus, we could observe the suggestive effect in column (4) and (5) of the upper panel of Table 2, though the it may still be marginal because the linear probability model does not support the statistical significance. Another possible interpretation is that childbearing and sometimes realization of homeownership transition arrive months after the decision is made. Two-year price change partly takes place after the family decision time point and thus has null explanation power. Nonetheless, a series of estimations on the marginal effect of lagged two-year price change all returns insignificant results, suggesting short-term price variation is lack of influence on general family behavior.

5.2 The other controls

The estimation results shown above suggest a local house price growth would cause a net negative effect on home buying but net zero effect on childbearing among households who chose not to move to another MSA. Of course, other controls also play an important role in the family decision. Table 4 reports the full table of marginal effects of controls less the fixed effects of two selected estimations, grouped in three column groups, which correspond to the model of column (1) and (4) of the upper panel of Table 2. Each column group consists three columns, in tandem reporting the marginal effects on entering homeownership only, giving a birth only, and doing both.

The estimates of the controls show the consistency between the two models.[25] Partnership (either by marriage or cohabitation) always has positive effects on both homeownership and childbearing. Whether it is newly formed matters only the homeownership. Total family income is positive related to home buying, but has a negative effect on childbearing at a smaller scale, presumably due to the greater opportunity cost of working time that leads to a net substitution for children. This substitution effect of working time is also reflected by the negative effect of the employment status of the woman, which leads to 1.3 to 1.4 percentage decrease in the probability of childbearing, in consistent with the classical fertility model (e.g. Becker, 1960).

Demographically, parenthood is a positive indicator for extra child, but the ability to enter homeownership may be deferred when the number of children is greater than two. This is probably because the big family size erodes the financial affordability of homeownership. Non-homeowning black women are much less likely to enter homeownership comparing with the white women, while they have a relatively higher likelihood to give a birth. It should be noted that this does not imply a disparity of fertility rate between races but a higher likelihood of parenthood without homeownership for black families. This difference may not only be attributable to the social and economic inequality but also the divergence in social norms on expected life course. Woman education shows a positive effect on both homeownership and childbearing. Again, we should be cautious of the interpretation as the baseline is the women with low education but already independent from her parents. The positive effect can reflect to the delayed fertility due to prolonged education time, so it is not necessarily reflecting the fertility difference by education. Regarding age, it seems the woman's age does not have a privilege or penalty on entering homeownership, but inevitably

[25]Not showing in the table, the estimates are also consistent between the model for two-year price change and four-year price change.

the childbearing likelihood declines over age steadily. All these observations are in line with the demographic regularities.

The MSA-level personal income per capita is the most important regional macroeconomic factors to the family decision. Higher personal income implies higher house price level in the local area. The unconditional correlation coefficient between the two variables in the sample is 0.71. Naturally as the result a high personal income level exerts the similar effect as the high house price level, and the effect is suppressed when both variables are included, as in column (1-1). However, this is not true for the change of personal income. By the estimations unshown here, short-term change of local personal income does not intervene family decisions on homeownership and childbearing directly. Recession has a negative impact on becoming a homeowner in that year, but in general has no substantial effect on the probability of childbearing.

So far, the estimates reflect the observed average marginal effect of house price level and changes. As Table 4 indicates that other factors are also associated with the probabilities of home buying and childbearing, it is reasonable to argue that the effect of house price level and changes are different for families under different financial and demographical conditions. Earlier study suggests non-linear and interaction effect are common in the decision of family homeownership (Li, 1977). Here, we briefly examine this potential heterogeneity of the effect by estimating an interaction term of the main variables of interest with three most outstanding controls, the real term family income, partnership, race, and parenthood. Without diving into this issue too deep, as each of the interaction effect is potentially a topic pending for further research, we look into interaction effects by showing the estimated marginal effects and the exponentiated coefficient of the interaction terms.

The results of the interaction effects with total family income are presented in Table 5. Panel A reports the ratio of relative-risk ratio (RRR) of the multiplicative term.[26] For the alternative of birth only and doing both, none of the estimates significantly stray from 1, indicating that family income level does not affect the relative volume of the effect. The estimates for the effect on entering homeownership are different. For house price level, the ratio of RRR is 0.87, meaning that the RRR in average is about 0.87 times for women with \$100,000 higher in total family income. Because the effect of house price on ownership is negative, the RRR of the effect is less than 1. As shown in Panel B, the RRR of house price level on entering homeownership is 0.75, which means the odds of entering homeownership would drop one quarter given a \$100,000 increase in house price level. The 0.87 times of

[26]Appendix A provides a short introduction to ratio of RRR.

Table 4: The MNL model estimates of the marginal effects of other controls.

Controls	(1-1)	(1-2)	(1-3)	(2-1)	(2-2)	(2-3)
	Dependent Choice					
	Ownership only	Birth only	Both	Ownership only	Birth only	Both
Median House Price ($100,000)	-0.046***	0.017***	-0.011***			
	(0.0143)	(0.0059)	(0.0039)			
4-Year Price Change ($100,000)				-0.016	0.004	-0.013***
				(0.0159)	(0.0085)	(0.0045)
Partnership	0.054***	0.036***	0.030***	0.054***	0.037***	0.030***
	(0.0184)	(0.0085)	(0.0096)	(0.0186)	(0.0085)	(0.0097)
Enter Partnership	0.040**	-0.002	-0.003	0.041**	-0.002	-0.003
	(0.0193)	(0.0107)	(0.0072)	(0.0193)	(0.0106)	(0.0074)
Total Family Income ($100,000)	0.083***	-0.028**	0.002	0.081***	-0.029**	0.001
	(0.0218)	(0.0123)	(0.0040)	(0.0221)	(0.0124)	(0.0038)
Employment	-0.010	-0.013*	0.009	-0.009	-0.014*	0.009
	(0.0212)	(0.0079)	(0.0065)	(0.0212)	(0.0080)	(0.0066)
Number of Children (base = 0)						
One	-0.001	0.058***	0.015**	0.000	0.058***	0.016**
	(0.0179)	(0.0083)	(0.0060)	(0.0180)	(0.0083)	(0.0062)
Two	0.004	0.086***	0.021***	0.005	0.086***	0.021***
	(0.0187)	(0.0119)	(0.0079)	(0.0189)	(0.0119)	(0.0079)
More than 2	-0.067***	0.124***	0.021**	-0.067***	0.123***	0.022**
	(0.0184)	(0.0142)	(0.0098)	(0.0184)	(0.0143)	(0.0098)
Race (base = other)						
White	-0.011	0.020*	0.001	-0.008	0.020*	0.001
	(0.0275)	(0.0115)	(0.0077)	(0.0280)	(0.0116)	(0.0078)
Black	-0.064**	0.027**	-0.004	-0.063**	0.028**	-0.005
	(0.0286)	(0.0129)	(0.0083)	(0.0289)	(0.0131)	(0.0082)
Education (base = no high school)						
High school diploma	0.043**	-0.000	0.004	0.044**	0.000	0.004
	(0.0186)	(0.0079)	(0.0032)	(0.0185)	(0.0079)	(0.0033)
Some college	0.030	0.021**	0.011**	0.032*	0.021**	0.011**
	(0.0185)	(0.0095)	(0.0050)	(0.0186)	(0.0095)	(0.0051)
College graduate	0.089***	0.021*	0.024***	0.088***	0.022*	0.024***
	(0.0218)	(0.0123)	(0.0084)	(0.0218)	(0.0125)	(0.0084)
Age group (base = 20-24)						
25-29	-0.015	-0.030	-0.012	-0.015	-0.028	-0.014
	(0.0401)	(0.0230)	(0.0190)	(0.0403)	(0.0228)	(0.0193)
30-34	-0.024	-0.063***	-0.026	-0.024	-0.063***	-0.028
	(0.0405)	(0.0224)	(0.0193)	(0.0406)	(0.0222)	(0.0197)
35-39	0.000	-0.095***	-0.035*	-0.001	-0.094***	-0.037*
	(0.0418)	(0.0222)	(0.0186)	(0.0420)	(0.0220)	(0.0189)
40-44	-0.027	-0.117***	-0.039**	-0.028	-0.116***	-0.041**
	(0.0427)	(0.0219)	(0.0188)	(0.0428)	(0.0217)	(0.0192)
State umemployment rate	-0.003	-0.001	-0.000	-0.004	-0.001	-0.003
	(0.0046)	(0.0023)	(0.0018)	(0.0052)	(0.0027)	(0.0021)
MSA personal income per capita	0.020	-0.057	0.135***	-0.365***	0.103*	0.055
	(0.1745)	(0.0770)	(0.0484)	(0.1173)	(0.0527)	(0.0406)
Average recession indicator	-0.050*	0.005	-0.009	-0.047	0.005	-0.012
	(0.0302)	(0.0139)	(0.0083)	(0.0304)	(0.0140)	(0.0089)
N	13026	13026	13026	13026	13026	13026

Note: Models differ only in the independent variable of interest. All estimates includes geographic (census division) and time (five-year period) fixed effects. Robust standard errors in parentheses account for clustering at individual level. Region (census division) and time (five-year) fixed effects in all specifications. Significant at *** p <10%, ** p <5%, * p <1%.
Source: Non-homeowning women who are either family head or its partner, age between 20 and 44, stayed in the same MSA in the last four years, with positive weight and all information available from PSID 1985-2015.

0.75 is about 0.65, indicating that, given a \$100,000 increase in total family income, the RRR of house price level is lower. In other words„ or the impact of house price on home buying is relatively larger for higher income group. Differently, the ratio of RRR of four-year price change is greater than 1, indicating its impact is smaller for women with higher family income, while the average marginal effect is insignificant. In sum, this result suggests that families with higher income are more responsive to high price level and more resilient to price change in home buying, but they have no significant difference in childbearing decision. These families have more financial capability against price appreciation.

It should be noted that this measure compares the RRRs, which is itself a ratio itself. The scale of RRR is determined by the marginal effect as well as the baseline odds. Because the baseline probability of home buying varies largely by income group, the ratio of RRR does not necessarily provide insights on the comparison of marginal effect in different income groups. To affirm the conclusion above, we show the marginal effects at different family income levels in Panel C. The marginal effects are consistent with the ratios of RRR. For house price level, the marginal effect is increasing with family income and for hour price change it is decreasing.

Ratio of RRR is not an intuitive measure. It is nevertheless a convenient tool to show the presence of interaction effect. Regarding partnership, race, and parenthood, the analysis remarks several notable points. We can see the results in Table A3. Women in partnership behave differently from who are not in partnership for home buying and doing both, though the big value of the latter results from the extremely small odds of doing both for women who are not in a relationship. There is also an interaction effect of two-year house price change and partnership on childbearing decision. Race difference creates a big divergence in the effect of two-year price change on childbearing decision, indicating different norms on children between black and white. Parenthood does not present a strong interaction effect except for doing both, which is again due to the extremely small case of doing both as the first birth. At any rate, this exercise shed some light on the more intricated mechanism of family decision. The more solid argument requires much more deeper investigations than the simple interaction term analysis.

5.3 Other Specifications

The estimated marginal effect might not reflect the true underlying mechanism if the statistics is unique to certain specifications. Sample selection, weighting, clustering, and other specification on the variable could all affect the estimated results. Here we test the

Table 5: The estimated ratio of RRR of the interaction effect of total family income and the marginal effects of house price level and variations, by income level.

		(1)	(2)	(3)	(4)	(5)
		Interaction with total family income				
A. Ratio of RRR	Dependent Choice	Median House Price ($100,000)	2-Year Price Change ($100,000)	2-Year Price Growth Rate	4-Year Price Change ($100,000)	4-Year Price Growth Rate
	Homeownership only	0.869*	1.582	3.081	1.613**	2.797*
		(0.0640)	(0.5295)	(3.2624)	(0.3839)	(1.7478)
	Birth only	1.104	1.378	1.211	1.266	1.785
		(0.1702)	(0.7253)	(1.8615)	(0.5208)	(1.8893)
	Both	1.003	0.768	0.211	1.160	1.170
		(0.1515)	(0.2798)	(0.3091)	(0.4796)	(2.0645)
	N	13026	13026	13026	13026	13026
B. RRR						
	Homeownership only	0.745***	0.495**	0.107***	0.559**	0.232**
		(0.1035)	(0.1636)	(0.1156)	(0.1276)	(0.1462)
	Birth only	1.237	0.657	0.785	0.892	0.89
		(0.1932)	(0.2697)	(1.0049)	(0.2405)	(0.7040)
	Both	0.419***	0.707	0.323	0.343*	0.068
		(0.1272)	(0.3969)	(0.6412)	(0.1898)	(0.1463)
C. M.E. on ownership only						
	Real Family Income at $10,000	-0.026**	-0.055**	-0.184**	-0.044**	-0.114**
		(0.0122)	(0.0275)	(0.0904)	(0.0194)	(0.0534)
	$30,000	-0.032**	-0.052**	-0.182**	-0.040**	-0.108**
		(0.0130)	(0.0262)	(0.0841)	(0.0178)	(0.0493)
	$50,000	-0.040***	-0.048*	-0.176**	-0.035**	-0.097**
		(0.0139)	(0.0257)	(0.0800)	(0.0168)	(0.0466)
	$70,000	-0.048***	-0.043	-0.166**	-0.027	-0.083*
		(0.0153)	(0.0269)	(0.0809)	(0.0171)	(0.0468)
	$90,000	-0.056***	-0.035	-0.151*	-0.017	-0.065
		(0.0172)	(0.0304)	(0.0894)	(0.0193)	(0.0515)

Note: Models differ only in the independent variable of interest. All estimates includes controls for partnership, race, number of children, total family income, and employment status, age group dummies, educational attainment dummies, state-by-year unemployment rate, MSA-by-year real income per capita, national recession index, and geographic (census division) and time (five-year period) fixed effects. Robust standard errors in parentheses account for clustering at individual level. For the panel of ratio of RRR, the value in parentheses reports the robust standard errors times ratio of RRR. Region (census division) and time (five-year) fixed effects in all specifications. Significant at *** $p<10\%$, ** $p<5\%$, * $p<1\%$.
Source: Non-homeowning women who are either family head or its partner, age between 20 and 44, stayed in the same MSA in the last four years, with positive weight and all information available from PSID 1985-2015.

robustness of our findings by estimating models with different specifications. Table 6 presents the test results with each panel reporting the estimates from a model one specification from the benchmark.

Panel A reports the results from the model that defines the alternatives as the actions taken place in the fixed two-year time window. As discussed in the empirical design, this setting is to prevent the uneven behavior accounting time after year 1997. The sample for this model only includes women from the odd year surveys, so that the sample sizes are noticeably smaller than in the benchmark model. The results are qualitatively similar to the upper panel of Table 2, but cannot reject the null hypothesis of the four-year marginal effect on doing both. This is not surprising. For the samples before year 1997, the new definition means a double length of the behavior time window. If a woman became a homeowner and have a child in two consecutive years before 1997, she is considered taking the two actions separately in each year the flexible time window scheme, but in the fixed time window scheme her behavior is classified as doing both during the two years period.[27] This could reduce the sensitivity of the suggestive impact of house price change.

The model for panel B uses the extended sample weight that includes women who join the survey because they enter the families of core survey members. Due to the reason, the newly added women have a much higher rate being in partnership and a higher average family income. This change accounts for an additional thirty-four hundred observations in the sample. At any rate, the estimates of the modified model are still consistent with the results in Table 2. Though the suggestive evidence of the marginal effect of four-year house price change is still marginal, expanding sample size does not really upset our main finding. All the reported standard errors so far are accounted for clustering individuals. As a panel data, it is reasonable because, under the framework of RUM, the unobservable components for the same individual over time are likely to be correlated. However, the geographical dimension of standard error correlation is also justified since our variables of interest and the controls for local economy are all region based. Thus, we re-estimate the standard errors by clustering samples by their residential MSA and present the results in panel C. Clearly, this modification does not change the main results. The standard errors are floating around the same level. Moreover, though not reported here, the combination of these specification change does not generate notable difference in the results.

[27]Because of that, the unconditional probability of doing both under the flexible time window is 0.6% and under the fixed time window is 1.1% for the same sample.

Table 6: The MNL model estimates of marginal effects of house price level and variations with other specification.

Dependent Choice	(1)	(2)	(3)	(4)	(5)
	Independent Variable				
A. Alternative dependent vairable: two-year fixed time window					
	Median House Price ($100,000)	2-Year Price Change ($100,000)	2-Year Price Growth Rate	4-Year Price Change ($100,000)	4-Year Price Growth Rate
Homeownership only	-0.053***	-0.014	-0.048	-0.005	-0.019
	(0.0152)	(0.0257)	(0.0782)	(0.0166)	(0.0461)
Birth only	0.021***	-0.007	0.006	0.007	0.030
	(0.0071)	(0.0135)	(0.0429)	(0.0095)	(0.0263)
Both	-0.016***	-0.005	-0.007	-0.008	-0.019
	(0.0057)	(0.0057)	(0.0332)	(0.0064)	(0.0200)
N	9074	9074	9074	9074	9074
B. Alternative weight: extended weight					
Homeownership only	-0.059***	-0.006	-0.050	-0.015	-0.056
	(0.0131)	(0.0221)	(0.0623)	(0.0141)	(0.0363)
Birth only	0.022***	-0.009	0.012	0.005	0.024
	(0.0061)	(0.0132)	(0.0339)	(0.0084)	(0.0203)
Both	-0.010**	-0.006	-0.023	-0.010*	-0.024
	(0.0045)	(0.0069)	(0.0226)	(0.0052)	(0.0150)
N	16340	16340	16340	16340	16340
C. Alternative clustering: clustering by MSA					
Homeownership only	-0.046***	-0.031	-0.137*	-0.016	-0.062
	(0.0137)	(0.0254)	(0.0748)	(0.0143)	(0.0389)
Birth only	0.017***	-0.007	0.011	0.004	0.021
	(0.0058)	(0.0109)	(0.0351)	(0.0097)	(0.0248)
Both	-0.011**	-0.007	-0.029	-0.013***	-0.034**
	(0.0043)	(0.0052)	(0.0195)	(0.0039)	(0.0145)
N	13026	13026	13026	13026	13026

Note: Models differ only in the independent variable of interest. All estimates includes controls for partnership, race, number of children, total family income, and employment status, age group dummies, educational attainment dummies, state-by-year unemployment rate, MSA-by-year real income per capita, national recession index, and geographic (census division) and time (five-year period) fixed effects. Robust standard errors in parentheses account for clustering at individual level. Region (census division) and time (five-year) fixed effects in all specifications. Significant at *** p <10%, ** p <5%, * p <1%.
Source: Non-homeowning women who are either family head or its partner, age between 20 and 44, stayed in the same MSA in the last four years, with positive weight and all information available from PSID 1985-2015. Panel A excludes all sample from the odd-year surveys. Panel B adds women with positive extended weight.

6 Discussion

How does house price affect the decision of non-homeowning families on home buying and childbearing? Our analysis presented in this paper helps us to sketch a big picture about the impact of house price level and variation for American urban families in the past thirty years. If it is not too arbitrary to assume these families have somewhat homogenous preferences on housing and children and are statistically representative, the results reveal a few key insights. First, a higher median house price would lower the probability of family entering homeownership and raise the probability of childbearing slightly, given other conditions unchanged. According to the statistics, a $100,000 increase in local median house price relates to a 5.7 percentage points net decline in the probability of becoming a homeowner and a 0.6 percentage points net increase in the probability of childbearing by and large. Unequivocally, home buying is sensitive to house price level for the obvious economic reason. Childbearing, differently, is affected by the high price not as much as housing. A marginal substitution for homeownership can only be inferred, not directly observed, from the estimations. More interesting is the dynamics of the two behavior. Women are more likely to give a birth without entering homeownership around the same time interval. In other words, more families decided to have a new child before become homeowner in areas with expensive median house price.

Second, the experience of two-year price change does not have an observable effect on the decision of home buying and childbearing. And the evidence reports only a weak negative effect of four-year price change on doing both in the same time window. This suggests the temporal change in house price does not have a strong impact on the family behavior, neither in absolute value nor ratio. There are two possible explanations. One is that recent local housing market variation simply does not alter family behavior on home buying and childbearing nor even their expectation on the future trend of equity value and child-rearing costs. Households care about only the current total cost of homeownership. The other is that such effects do exist, but the negative impact of lower relative income due to the increased price is offset by the positive expectation on future equity appreciation. We cannot directly tell which one is closer to the reality, but the negative marginal effect of four-year price change on concurrent home buying and childbearing hints that family's willingness to take the costliest move is eroded by house price growth, implying that the negative impact might surpass upon a high cost condition. Therefore, the argument of the co-existing offsetting effects is more plausible.

Third, no matter whether the average marginal effect is significant from zero, the results does not imply a linear effect across the whole sample. Families with higher income are hit more by high house price level to enter homeownership probably due to their higher unconditional likelihood of home buying. The effect of price appreciation behaves oppositely. Women with lower income are affected more by it, likely because they faced a tighter credit constraint and have a lesser chance to acquire benefit from equity appreciation. The analysis also shows significant interaction effects of house price with partnership and race, while whether the parenthood of women seems less critical. This finding signifies the complexity of family decision as the influence of a single factor is multi-dimensional, entangled with numerous other considerations. Greater economic inequality and declining marriage rate (but compensated by growing cohabitation rate) are both likely to play a role at the aggregated level (Lutz et al., 2006; Lesthaeghe, 2010). A more detailed mechanisms may hide beneath the surface, though it is out of the scope of this paper.

Considering the housing market only, house price may not directly affect current regional birth rate according to our results, but it not at all unimportant to family fertility. An expensive housing market would alter the family life course plan, push homeownership behind parenthood, letting more children be raised in rental unit during their infancy. Although in this paper only the homeownership is referred regarding family housing choice, it encapsulates the common differences between rental and owner-occupying housing units, including floor space, maintenance quality, tenure stability, surrounding amenity and facility, all of which could lead to a profound legacy to children, as Haurin and other authors argue. On the other hand, the recent lowering fertility rate nationwide seems not to be attributable to the rising house prices. At least for non-homeowners staying in the same city, growing house prices may only generate a temporary discouragement on childbearing for women with partners. Other economic and demographic transformation inside the society should have a greater and perpetual influence on aggregate fertility rate.

Our findings are in line with the literature and contribute to a deeper understanding of the association between the housing market and family homeownership and childbearing. The dynamics between the two family behavior is shown to be sensitive to the market variation. Nevertheless, this analysis has clear limitations. In order to prevent house price endogeneity, people who migrated to other metropolitan areas are excluded from our research. But migration is a crucial dimension in family life course. It allows family to actively choose the house price it would encounter and closely relates to family income and the condition of living environment. Though it is a relatively small group, empirically women from the migration group have a higher probability of childbearing, suggesting the importance of migration on fertility. Inversely, local house price variations or even spatially relative house price disparity can also change the migration decision and in tandem affect home buying and childbearing decision. In addition, the interaction of housing and childbearing is also influenced by other major life course transition such as partnership and employment, which are treated as exogenous in our analysis for the purpose of our research. They aren't. Regarding family formation, they are as substantial as housing and childbearing. Investigations on multi-dimensional choice model in a dynamic framework could reveal more insights to individual decisions, and this research is just a start. As we show the dynamics of major family transitions is sensitive to house price, it may well happen to partnership and career path. We look forward to more detailed researches to disentangle the underlying secret of the economic-demographic interplay.

Another challenge to the analysis of the impact of house price is the difficulty of accurately measuring the real cost families are facing. Besides the fact that house prices may vary in a remarkable range in big city and families have divergent housing demand, other factors, including the loan-to-value ratio, mortgage interest rate, and current rental cost, are also accountable for estimating the financial cost of homeownership. The credit constraint of home mortgage is in specific the major obstacle to homeownership, and its volume depends on the proportion of the property value that banks are willing to loan out. The mortgage interest rate also plays an important role as it determines the overall property cost. Unfortunately, we do not have the complete information about what kind of mortgage offer respondents can obtain. In this paper we instead assume the financial burden is exclusively proportion to the local median house price. We expect more questions about the joint family behavior could be answered with the help of a more detailed date of the real cost of homeownership in the future.

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Table A1: The IIA property test for the MNL model.

Independent Variable	Alternative Dropped	Null Hypothesis					
		$\hat{y}_1$ (full) = $\hat{y}_1$		$\hat{y}_2$ (full) = $\hat{y}_2$		$\hat{y}_3$ (full) = $\hat{y}_3$	
		χ^2	p-value	χ^2	p-value	χ^2	p-value
Median House Price ($100,000)	Ownership only			0.174	0.676	0.049	0.825
	Birth only	0.632	0.427			0.061	0.805
	Both	3.260	0.071	0.034	0.854		
2-Year Price Change ($100,000)	Ownership only			0.358	0.549	4.129	0.042
	Birth only	0.134	0.714			0.023	0.880
	Both	0.549	0.459	0.860	0.354		
2-Year Price Growth Rate	Ownership only			2.622	0.105	6.870	0.009
	Birth only	0.017	0.895			0.044	0.834
	Both	0.022	0.881	0.439	0.508		
4-Year Price Change ($100,000)	Ownership only			0.142	0.706	1.319	0.251
	Birth only	1.907	0.167			0.394	0.530
	Both	0.528	0.468	1.994	0.158		
4-Year Price Growth Rate	Ownership only			0.457	0.499	0.936	0.333
	Birth only	1.078	0.299			0.182	0.669
	Both	0.887	0.346	1.523	0.217		

Note: The null hypothesis for all tests is that the estimated odds of the alternative from the benchmark model (full alternatives) is the same with the estimated odds of the alternative from the model with one other alternative dropped. $\hat{y}_1$ denotes the odds "ownership only," $\hat{y}_2$ the odds of "birth only," and $\hat{y}_3$ the odds of "doing both."

Table A2: The MNL model estimates of elasticities of house price level and variations.

	(1)	(2)	(3)	(4)	(5)
Dependent Choice	Independent Variable				
No inter-MSA move in the past four years					
	Median House Price ($100,000)	2-Year Price Change ($100,000)	2-Year Price Growth Rate	4-Year Price Change ($100,000)	4-Year Price Growth Rate
Homeownership only	-0.715***	-0.014	-0.032*	-0.012	-0.025
	(0.2171)	(0.0112)	(0.0175)	(0.0098)	(0.0162)
Birth only	0.639***	-0.009	0.003	0.004	0.015
	(0.2350)	(0.0138)	(0.0218)	(0.0131)	(0.0221)
Both	-1.567***	-0.030	-0.064	-0.073***	-0.120**
	(0.5357)	(0.0227)	(0.0432)	(0.0235)	(0.0525)
N	13026	13026	13026	13026	13026

Note: Models differ only in the independent variable of interest. All estimates includes geographic (census division) and time (five-year period) fixed effects. Robust standard errors in parentheses account for clustering at individual level. Region (census division) and time (five-year) fixed effects in all specifications. Significant at *** p <10%, ** p <5%, * p <1%.
Source: Non-homeowning women who are either family head or its partner, age between 20 and 44, stayed in the same MSA in the last four years, with positive weight and all information available from PSID 1985-2015.

Appendices

A Ratio of RRR

Because our model is non-linear, the interaction effect cannot be simply identified by the coefficient on the interaction terms alone. Instead, the exponentiated coefficient of the multiplicative term between two explanatory variables can imply the presence of interaction effects (Buis, 2010). For a MNL model, the exponentiation of coefficient β_{ik} is called the RRR for alternative i of an independent variable x_k. It is defined as the ratio of the relative probability of i for a one unit increase in x_k. If the value is greater than one, it means that the relative probability of i is greater given an increase in x_k. This interpretation is derived from Equation (3).[28]

$$RRR(\beta_{ik}) \equiv e^{\beta_{ik}} = \frac{e^{x'\beta_i+\beta_{ik}}}{e^{x'\beta_i}} = \frac{P_i(x_k+1)/P_0(x_k+1)}{P_i(x_k)/P_0(x_k)} \tag{6}$$

The exponentiation of a multiplicative term is the ratio of RRR for the two explanatory variables (Norton et al., 2004). It tells the relative volume of effect, in term of RRR of one variable, for a one unit increase in the other variable. If we add an interaction term of x_k and x_l to the RUM model and let β_{ikl} be its coefficient for alternative i, we have

$$e^{\beta_{ikl}} = \frac{e^{x'\beta_i+\beta_{ik}+\beta_{il}+\beta_{ikl}}/e^{x'\beta_i+\beta_{il}}}{e^{x'\beta_i+\beta_{ik}}/e^{x'\beta_i}} = \frac{RRR(\beta_{ik} \mid x_l+1)}{RRR(\beta_{ik} \mid x_l)} \tag{7}$$

In our case, we set the first variable as the variable of interest, and the second one is the interacted control variable. The ratio of RRR shows how many times the RRR of the variable of interest would change given a unit increase of the control. If RRR is greater than 1, a greater ratio of RRR indicates the effect is intensified by the interaction. If less than 1, a greater ratio indicates the effect is diminished by interaction.

[28]Mathematically, RRR is also the proportion of the risk ratio of alternative i for a unit increase in x_k to the risk ratio of the base outcome for a unit increase in x_k. However, this form does not provide an intuitive interpretation.

Table A3: The estimated ratio of RRR of the interaction effect of partnership, race, and parenthood.

Interacted with:	Choice	(1)	(2)	(3)	(4)	(5)
		Independent Variable				
B. Partnership		Median House Price ($100,000)	2-Year Price Change ($100,000)	2-Year Price Growth Rate	4-Year Price Change ($100,000)	4-Year Price Growth Rate
	Ownership	0.811*	0.708	0.398	0.911	0.768
		(0.1020)	(0.2834)	(0.4484)	(0.2129)	(0.4580)
	Give a birth	0.955	0.346***	0.050**	0.869	0.501
		(0.1234)	(0.1315)	(0.0640)	(0.2143)	(0.3431)
	Both	36.587***	0.815	0.582	0.922	1.924
		(37.0115)	(0.3086)	(0.9397)	(0.3294)	(2.5156)
	N	13026	13026	13026	13026	13026
C. Black						
	Ownership	1.105	1.176	2.157	1.149	2.006
		(0.1650)	(0.5247)	(2.7851)	(0.3683)	(1.5135)
	Give a birth	1.104	2.609**	17.045*	1.328	3.183
		(0.1380)	(1.1921)	(28.0144)	(0.3847)	(2.7000)
	Both	0.333	2.782	43.557	1.263	2.610
		(0.2600)	(2.3649)	(146.8183)	(0.7174)	(4.7360)
	N	12359	12359	12359	12359	12359
D. Parenthood						
	Ownership	1.146	1.036	1.357	1.048	1.604
		(0.1497)	(0.3673)	(1.3729)	(0.2236)	(0.8876)
	Give a birth	1.088	0.784	0.816	0.602	0.275
		(0.1895)	(0.4504)	(1.3796)	(0.1981)	(0.2182)
	Both	1.065	2.851*	6350.399***	2.134	1857.790***
		(0.3561)	(1.6900)	(20222.2860)	(1.1907)	(4988.5000)
	N	13026	13026	13026	13026	13026

Note: Models differ only in the independent variable of interest. All estimates includes controls for partnership, race, number of children, total family income, and employment status, age group dummies, educational attainment dummies, state-by-year unemployment rate, MSA-by-year real income per capita, national recession index, and geographic (census division) and time (five-year period) fixed effects. Robust standard errors in parentheses account for clustering at individual level. For the panel of ratio of RRR, the value in parentheses reports the robust standard errors times ratio of RRR. Region (census division) and time (five-year) fixed effects in all specifications. Significant at *** p <10%, ** p <5%, * p <1%.
Source: Non-homeowning women who are either family head or its partner, age between 20 and 44, stayed in the same MSA in the last four years, with positive weight and all information available from PSID 1985-2015.