

Homes and Husbands for All: Marriage, Housing, and the Baby Boom

Matthew Hill

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Abstract

The U.S. experienced an unprecedented increase in fertility during the baby boom. After falling for a century and a half, fertility rates jumped by 45 percent from 1945 to 1955, before resuming their decline in 1965. The elevated birth rates from 1946 to 1964 were driven in part by a shift toward more universal marriage; marriage rates increased by 25 percent from 1930 to 1950 and the average age of marriage fell by two years. This paper argues that growth in the supply of housing after World War II contributed to the expansion of marriage during this period. Specifically, the paper estimates the effect of additional building permits (a proxy for housing supply) at the city level on individual marriage outcomes. An instrumental variable approach is used to address endogenous permit location. I construct an annual level instrument using the national permit series in conjunction with a city's geographical constraints, region and average temperature. I find a standard deviation increase in permits to a city increased the probability of marriage in that city by 13 to 16 percent over a two-year period. The estimates suggest that the growth in housing supply in the late 1940s can explain about 33 percent of the difference in marriage rates between 1930 and 1950. Overall, the increase in housing supply can account for nearly ten percent of the baby boom.

1 Introduction

During the baby boom (1946-1964) an average of 4.2 million babies were born per year, compared to 2.4 million per year in the 1930s. From 1946 to 1964, annual fertility rates were 30 to 50 percent higher than they were in the pre-war era. This increase in fertility was unprecedented in the United States. Fertility

rates had been falling by 8 percent per decade between 1800 and 1940 and declined by 10 percent per decade between 1965 and 2005. The rise in birth rates in the baby boom is a result of two trends: first, there was an increase in birth rates within marriage and second, marriage itself became more universal. Marriage rates increased by 25 percent from 1930 to 1950, and the average age of marriage fell by two years.

Explanations for the baby boom have focused on the increase in birth rates within marriage. The primary hypotheses for this rise in fertility are that young couples who had come of age during the Great Depression had a preference for children over material goods (Easterlin), that improvements in household technology lowered the cost of children (Greenwood et al.), and that declines in maternal mortality led to an initial increase in the quantity of children (Albanesi and Olivetti). In contrast to the previous literature, I examine the cause of the rising marriage rates.

I argue that a burst in new housing construction, after a near construction freeze during the Great Depression (1929-1940) and World War II (1941-1945), can account for 33 percent of the observed change in marriage patterns from 1930 to 1950. After World War II, new residential housing boomed; 1.3 million residential units were built in 1950, over triple the number constructed in any depression-era year. However, the magnitude of this housing boom varied from city to city. I use a newly digitized dataset on permits per capita as a proxy for housing supply and exploit the variation between cities to estimate the effect of new permits per capita on marriage outcomes. I find that a standard deviation increase in permits per capita in a given year raised the probability of marriage by 6 to 8 percent in the subsequent year.

In theory, couples in areas where more homes are built should marry sooner, *ceteris paribus*, than couples in areas with fewer homes because of the long-standing cultural norm in Western societies that couples obtain their own housing upon marriage. Even in the pre-industrial age, the majority of Northern European households were made up of nuclear families (Hajnal, 1962). By the mid-twentieth century, this norm had not changed; one married couple per household remained the standard arrangement. For instance, in 1940, 93 percent of married twenty-five year-olds had their own household while just 44 percent of their non-married peers lived on their own. Given the expectation of setting up a household upon marriage, lower housing prices, due for example to increased housing supply, should raise marriage rates and speed transitions into marriage.

A potential concern in estimating the effect of permits per capita on marriage is that new building permits may be a response to marriage. Builders may observe marriage rates rising and build homes to accommodate the rush of new couples. Thus, a positive correlation between building permits and marriage outcomes in the cross section may be driven by construction spurred by new marriages rather than the effect of new construction on marriage. I employ a fixed effects specification that should mitigate the bias from reverse causality because in this specification marriage outcomes in a two-year period are paired with permits per capita in the previous two-year period. However, if builders are forward looking, the bias from reverse causality will still be present in this specification.

A second potential concern is that omitted variables may bias the estimates of the effect of permits per capita on marriage outcomes. For instance, a local labor demand shock might drive new construction as builders respond to recent migrants. The same demand shock also might increase entry into marriage because new jobs and/or increased job security provides the economic stability that facilitates marriage. To address the potential bias of the estimated effect of permits per capita on marriage rates caused by omitted variables and reverse causality, I construct an instrument for annual permits. The instrument is built by allotting aggregate annual permits to the sample cities based on two weights that are unrelated to contemporaneous labor demand shocks. The first weight is a geographic weight: cities that are more geographically constricted receive fewer permits than less constricted cities. For example, if a city can only build on half of its land because of geographic features like marshland or steep slopes, it receives a geographic weight of .5. The second weight is a population weight. A population growth weight is the appropriate weight because permits represent the flow of new housing construction, rather than the existing stock. However, population growth may also be correlated with omitted variables, such as local labor demand shocks. Thus, ten-year lagged predicted population growth is used as the weight, where population growth is predicted by region and average annual temperature. Essentially, I use three time-invariant factors (region, geography, and temperature) unrelated to local labor demand shocks to generate an annual instrument that exploits national variation in housing demand.

This paper contributes not only to the baby boom literature but also to the emerging literature on the effect of housing supply. Recent work suggests several

cities have increasingly restricted their housing supply through regulation.¹ The immediate effect is that the cost of housing has risen dramatically in these areas over the past 20 years. An artificially restricted housing supply has indirect effects as well: Glaeser, Gyourko and Saiz (2008) show restricted supply leads to more volatile prices; also higher housing costs may drive away firms that are unwilling to pay higher wages. This paper shows that restricting housing supply will lead to lower marriage rates and higher ages at first marriage. Cities with more housing regulations already have lower rates of in-migration; my results suggest that these cities will also have lower levels of fertility as well.

The paper proceeds as follows: Section 2 provides background on the expansion of marriage in the period and the state of the postwar housing markets. Section 3 outlines the conceptual framework for the effect of housing costs on marriage rates. Section 4 details the datasets used in the empirical estimation. Section 5 presents the instrumental variable analysis used in the paper and the results of the estimation. Section 6 discusses the magnitude of the estimates and the role they played in the baby boom. Section 7 concludes.

2 Background on the 1950s Baby Boom and Housing Market

2.1 Marriage and the Baby Boom

From 1930 to 1950, marriage rates increased by 25 percent and the average age at first marriage for both men and women fell by two years. These marriage trends have significant fertility implications because married women have higher birth rates than single women. The difference in birth rates between single and married women is even more pronounced in the baby boom period because single motherhood was a rare event.² Furthermore, contraception was less reliable in the 1950s than the present.³ Thus, when women married younger, they were likely to have more children because they faced a higher risk of pregnancy for more years.

How large a role did the marriage trends of the period play in the baby boom? To connect the expansion of marriage to increased birth rates more

¹See Glaeser, Gyourko, and Saks (2005), Saiz (2010), and Glaeser and Ward (2009)

²The birth rate for single women was 6 times higher in 1998 than it was in 1950. (Carter et al. 2006)

³Oral contraceptives were not introduced until 1960.

formally, I present a decomposition of the change in the crude birth rates from 1930 to 1950. This decomposition shows the share of the aggregate increase in crude birth rates that can be explained by the changes in marriage patterns. I decompose the difference in crude birth rates as follows:

$$cbr_{50} - cbr_{30} = \sum_{am=54}^{AM} (cbr_{am50} - cbr_{am30})\rho_{am50} + \sum_{am=54}^{AM} (\rho_{am50} - \rho_{am30})cbr_{am30}$$

where am indexes each distinct demographic group, and groups are defined by age and marital status. Ages 14 through 40 are included in the decomposition, yielding 54 total groups. The variable, ρ_{amy} , is the share of individuals in demographic group am in year y . The first term gives the difference in crude birth rates attributable to rising birth rates within groups. The second term measures the difference in crude birth rates due to a shift in the proportions of the population in each demographic group, i.e., the difference in crude birth rates due to increases in the proportions married at given ages. The total difference in crude birth rates between 1930 and 1950 is 35 births per 1000 women aged 14-40. The calculation yields a value of 25 for the first term and 10 for the second term. Thus, marriage patterns alone account for 29 percent of the difference in crude birth rates between 1930 and 1950. If the calculation is done for 1960, marriage patterns account for about 25 percent of the difference in crude birth rates between 1930 and 1960.

Very few papers tackle the cause of the marriage boom in the postwar period. Doepke, Hazan, and Maoz (2007) argue that young women after World War II faced stiff labor force competition from both older women who had been employed during the war effort and returning male veterans. The authors posit that with few job opportunities, young women were then driven to marriage. This hypothesis is inconsistent with two facts. Firstly, Albanesi and Olivetti (2009) show that the labor force participation of mothers actually rose in the period. If women were driven to marriage because of lack of jobs, why did their labor force participation increase? Secondly, the Doepke, Hazan and Maoz hypothesis suggests that young women should face stiffer labor force competition in states with higher war mobilization rates and be driven to marry earlier than in states with lower war mobilization rates. However, states with the highest rates of war mobilization had the highest ages of marriages for the cohort of women most affected by labor force competition after World War II.⁴ Other

⁴I define the cohort most affected to be women who were in high school during World War

papers not directly related to the marriage boom offer clues as to why marriage increased in the postwar period. Hill (2011) finds evidence that marriage rates were sensitive to GDP prior to 1960, and Loughran (2002) shows that women’s propensity to marry increases when the wage distribution of men is more equal. The 1950s was both an era of prosperity and the period in U.S. history when the wage distribution was most equal. These two factors likely played a role in the increase of marriage rates after World War II; however, there may be other trends that also positively influenced marriage rates, such as the growth in the housing supply.

2.2 Housing and Building Permits

In 1945, on the eve of the baby boom, the U.S. faced a housing shortage. A report to Congress estimated that the country needed 3,000,000 homes to accommodate the returning veterans and their prospective families.⁵ The Great Depression and World War II had dampened construction for the previous 15 years. A Senate report scolded, “The lack of decent housing within economic reach of all American families may once have been a national scandal. It is now a national tragedy.”⁶ The U.S. government was concerned enough about the dearth of housing that it formed a new agency, “The Office of the Housing Expediter,” to help handle the crisis.

On the housing demand side, the number of people who could afford to purchase a home after World War II increased. In 1944, Congress passed the G.I. Bill, officially the Servicemen’s Readjustment Act, which included several provisions specifically related to housing. Through the Veterans Administration mortgage insurance program, veterans were granted access to home loans with low interest rates and very low down payments.⁷ The housing provisions had widespread uptake; from 1944 to 1952, 2.4 million veterans received a federally insured home loan and about 35 percent of all home mortgages were federally insured. Fetter (2010) estimates that 70 percent of the increase in homeownership from 1940 to 1960 among 26 year-olds and 32 percent of the increase for 32 year-olds can be attributed to G.I. Bill benefits. The G.I. Bill’s influence went beyond guaranteeing loans; it changed the features of mortgages by proving to

II and therefore too young to have worked during World War II.

⁵Congress, House, Committee of the Whole, Director of Housing Stabilization. 79th Cong., 2nd Sess. Report No. 1580.

⁶Congress, Senate, Committee on Banking and Currency, Veterans’ Emergency Housing Act of 1946. 79th Cong., 2nd Sess. Report No. 1130.

⁷see Green and Wachter (2005)

private lenders that loans with extended terms and low down payments were economically feasible. By 1947, home loans had a very different structure than the home loans of twenty years prior. Compared to 1920, the average loan in 1947 was almost twice as long, had a 30 percent lower interest rate, and a 30 percent smaller down payment. These changes in home loans can be thought of as an advancement in mortgage technology that lowered the overall cost of housing. The types of mortgages propagated after World War II are especially relevant in the context of the baby boom because home loans with less stringent savings and credit constraints enabled couples to transition into marriage at earlier ages.

I argue that permits are a good proxy for the cost of housing. Permits per capita will capture the differences in supply elasticity between cities. In the postwar period, demand for housing was shifting upward as all cities increased in size and the GI provisions lowered the cost of housing for buyers. Builders responded to this demand shift. Large scale developers, like Levitt, constructed whole new cities and enterprising manufacturers, like Lustron, provided mail order pre-fabricated houses. Figure 1, a plot of building permits per capita from 1933 to 1958, shows new permits per capita increased six fold from 1945 to 1950. However, cities varied in their ability to respond to the increased demand for housing. Places with ample land and/or the presence of entrepreneurial developers would have been able to rapidly increase their housing stock, while more constricted, low supply elasticity cities would have been slower to respond. In these more constricted cities, home prices would have been higher, as sellers could extract a premium from buyers. Table 1, a regression of home prices on permits, shows that this was the case. There is a strong negative correlation between new housing permits from 1944-1949 and home prices in 1950. A 60 percent increase in permits per capita (the standard deviation between cities) is associated with a six percent decrease in prices.

3 Theoretical Framework

In this section, I present a simple model to clarify the relationship between marriage and housing. Consider a three-period model representing a woman's early, mid, and late twenties. Women are assumed to be similar, thus assortative mating is not a component of the model as it is in Burdett and Coles (1997) and Bloch and Ryder (2000). This is done because I am not interested in who marries

who rather I am concerned about the timing of marriage. This decision is also informed by the historical context: there was much less assortative mating in the 1950s than the present (Schwartz and Mare 2005).⁸ In each period in which a woman is single, she receives a draw from a distribution of potential mates. If she accepts the draw she remains married to the mate for all periods thereafter, i.e., there is no divorce. The goal of the model is to find an expression for the reservation value in period t (r_t) above which she will accept her mate draw, and to examine the effects of housing costs on r_t . If she rejects the mate, a woman receives zero utility for that period. If she accepts the mate, she gets utility equal to $(1 - \phi(t))x_t$.⁹ Where x_t is the value of her draw from period t and $\phi(t)$ represents the cost of housing. Let $\phi(t)$ be decreasing in t , reflecting the fact that housing is less costly at older ages because credit is more readily available and older men have more savings. For simplicity, I assume $\phi(1) = c$, $\phi(2) = \frac{c}{2}$, $\phi(3) = 0$. Intuitively, women weigh the trade-off between accepting a mate in the current period and holding out for a better option. Housing costs act as a tax on current consumption so women are more likely to wait when costs are high. It is straightforward to show the reservation value in period 3, r_3 , is zero. In period 1, if the woman accepts her draw, her utility is:

$$U^{Accept} = (1 - c)x_1 + \beta(1 - \frac{c}{2})x_1 + \beta^2x_1$$

If she declines, her utility will be:

$$U^{Decline} = 0 + Pr(x_2 > r_2) \left(\beta(1 - \frac{c}{2})E[x_2|x_2 > r_2] + \beta^2E[x_2|x_2 > r_2] \right) +$$

$$Pr(x_2 < r_2) \left(\beta^2Pr(x_3 > 0)E[x_3|x_3 > 0] \right)$$

If she declines in the first period, she will accept in period 2 if

$$\left(x_2(1 - \frac{c}{2}) + \beta x_2 \right) > (\beta Pr(x_3 > 0)E[x_3|x_3 > 0])$$

From the above equations, one can show:

⁸Also the late 1940s had a much lower skill premium than the present (Juhn 1999) and Fernandez, et al. (2005) show assortative mating decreases when the skill premium falls.

⁹In the simulations I will assume the distribution of males is normal, with mean 0 and variance 1. This is done for ease of computation. The implications of the model are valid for any distribution of males.

$$\begin{aligned}
r_1 &= \left(\frac{1}{1 - c + \beta(1 - \frac{c}{2}) + \beta^2} \right) (Pr(x_2 > r_2) \left(\beta(1 - \frac{c}{2})E[x_2|x_2 > r_2] + \beta^2E[x_2|x_2 > r_2] \right) \\
&\quad + Pr(x_2 < r_2) (\beta^2 Pr(x_3 > 0)E[x_3|x_3 > 0])) \\
r_2 &= \left(\frac{\beta}{\beta + 1 - \frac{c}{2}} \right) Pr(x_3 > 0)E[x_3|x_3 > 0]
\end{aligned}$$

And

$$\frac{dr_1}{dc} > \frac{dr_2}{dc} > \frac{dr_3}{dc} = 0$$

Suggesting that as housing costs (c) rise, so does the reservation value in periods 1 and 2. Figure 2, a simulation of the model, illustrates the difference in marriage rates under different housing cost regimes. When housing costs are low ($c=.1$), 37 percent of women will marry in period 1 and by period 2, 64 percent of women will have married. When housing costs are high, 31 percent of women will marry in period 1 and 59 percent of women will have married by period 2.

The model has two other key implications. Housing costs affect the reservation wage more in earlier periods, implying that the estimated effect of housing on marriage should be larger for younger age cohorts than for older age cohorts. A second implication is that housing costs should have a larger impact on men with lower x values. In figure 2, the difference in marriage rates between the low housing cost regime and the high housing cost regime is driven by the willingness of women in the low cost regime to accept a man with a lower x_1 value. Men from the higher end of the distribution are unaffected by housing costs; they marry under both regimes. Empirically, one should observe that housing costs have larger effects for men on the lower end of the distribution.

4 Data

Information on marriage outcomes is available at both the individual and county level. The individual-level data I use is from the 1950 census. This dataset contains an individual's current marital status and the duration of that status. From this marriage data, I can ascertain whether an individual was single or

married in a given year and therefore construct a retrospective marriage history for each individual. The census data also includes an individual's current city and whether the person lived in this city one year prior. The location information allows me to merge individual data with data on the relevant city from other sources. Additionally, the census has background characteristics, including race, education, age, labor-force status, and veteran status. A drawback of the census data is that there may be some migration bias introduced because I assume an individual lived in the same location from 1946 to 1950. However, in the empirical section I report estimates for a sample with less migration bias and show that my results do not change significantly.

After 1950, I need to rely on county-level marriage data because the 1960 Census microdata do not report an individual's location below the level of the state. I use marriage licenses issued in a county as the marriage outcome for the second half of the baby boom period: 1954-1962. This data was available for 66 of the most populous U.S. counties from the U.S. National Office of Vital Statistics' Monthly Vital Statistics Reports.

Building permit data come from the Bureau of Labor Statistics (BLS). Beginning in 1920, the BLS began collecting data on new permits issued in urban areas. The dataset is at the city-level and only includes buildings undertaken within a city's limits. However, new permits within a city are highly correlated with new permits in the surrounding areas (correlation coefficient of 0.74 for the years when data were available, 1962-64). In 1954, the permit data series was taken over by the Census Bureau. The Census Bureau changed the level of observation from the city-level to the county-level by collecting data on towns and unincorporated areas surrounding a central city. I digitized the city-level permit data from 1921 to 1954 and the county level permit data from 1954 to 1964 for use in this paper.

Data on other city and county characteristics are from the 1944, 1948, and 1952 city data books and the 1956 and 1962 county data books. This dataset includes other economic characteristics that might influence the marriage decision, including employed persons and retail sales per capita (a proxy for local GDP). I merge the city data books with the individual-level data from the 1950 census. I chose city-level data to be consistent with the permit data from this time period. I merge the county data books with the post-1954 marriage license and building permit data, which are both also observed at the county-level.

5 The Effect of Housing Supply on Marriage

5.1 Empirical Design

I examine the effect of the number of new building permits per capita in an individual's city on their marital status in 1950. I estimate a linear probability model for a cross section in 1950.¹⁰ The estimating equation is:

$$Marriage_{ij} = \alpha + \beta \cdot PerCapPermits_j + y_j' \phi + x_i' \delta + \varepsilon_{ij} \quad (1)$$

where $Marriage_{ij}$ is a dummy variable equal to 1 if individual i in city j married in the previous three years; it equals 0 if the individual was single in 1947 and remained single through 1950. $PerCapPermits_j$ is the per capita permits in city j from 1947 to 1949, y_j is a vector of city controls including employment, sex ratio and retail sales, and x_i is a vector of individual controls including age, age squared and an education dummy equal to 1 if the individual has 12 or more years of schooling. The sample is composed of white males age 20-27 and white females age 18-25 who were single in 1947. I limit the sample to these particular age groups because they are the prime marriage ages for the respective sexes in the period, by age 25, 78 percent of women were married and by age 27, 72 percent of men were married. The results of the estimation are robust to altering the sample to include other ages around the prime marriage years.

I also employ a city fixed effects specification that has two advantages over the 1950 cross section. Firstly, it controls for time-invariant factors that may influence both marriage outcomes and building permits that are not accounted for by the city controls in the cross-section specification. For instance, politically conservative cities may have more construction due to a more permissive regulatory environment, and they may also have a culture of early marriage due to widely held religious beliefs. Secondly, in this specification I relate contemporary marriage outcomes to lagged permits because there is an inherit lag in the permitting-to-construction process. It takes about one year for a permit to become a house. The use of lags also mitigates issues with reverse causality that stem from construction induced by rising marriage rates. In the fixed-effects specification, the dependent variable is a dummy equal to 1 if the individual married in a two-year period. I use two-year time periods to mitigate measurement error introduced by calculating a marriage year from the census variable:

¹⁰Results are robust to using a probit or logit specification.

“duration of current marital status.” As in the cross section, I limit the sample to individuals who were white, single and age 20-27 (18-25 for females) at the start of each period. The estimating equation is:

$$Marriage_{ijt} = \alpha + \gamma_j + \chi_t + \beta \cdot PerCapPermits_{jt} + x_i' \delta + \epsilon_{ijt} \quad (2)$$

where $Marriage_{ijt}$ is a dummy variable equal to 1 if individual i in city j married during period t . $PerCapPermits_{jt}$ is per capita permits from the previous two-year period, for instance if the marriage period is 1948-1949, then permits from 1946-1947 are used. The term, γ_j is a city fixed effect and χ_t is a period fixed effect.

For the second half of the baby boom period when individual data is not available, I perform the empirical analysis at the county level. The estimating equation is:

$$PerML_{jt} = \alpha + \chi_j + \beta \cdot PerCapBP_{jt} + \varepsilon_{jt} \quad (3)$$

$PerML_{jt}$ is the marriage licenses per capita in county j in year t , $PerCapBP_{jt}$ is the building permits per capita in county j in year t , and χ_j is a county fixed effect. Here the sample is 66 counties where the data are available from the year 1954 to 1962.

5.2 Instrumenting for Building Permits

Estimates of the effect of building permits on marriage outcomes may be biased by three potential sources: reverse causality, omitted variables, and measurement error. As discussed in 5.1, the fixed effect specification addresses the issue of reverse causality. However, estimates from the fixed effect specification may still suffer from omitted variable bias. The city fixed effect will control for the long run economic prosperity of cities but will not account for the presence of a one-period labor demand shock. For instance, the construction of a new factory could influence both marriage outcomes and new building. Finally, measurement error may be present in the building permit data. The Bureau of Labor Statistics only collected permits within a city’s limits. Therefore, building outside a city is not included in the data.

To address the potential biases in the estimates, I generate an instrument called predicted permits. The instrument is constructed by allocating aggregate

permits to the sample cities based on geographical conditions unrelated to contemporaneous demand shocks. Essentially, I build an annual instrument from a city’s time-invariant factors by using variation in the national permit series. The national permits are assigned to cities based a on a city-specific weight. The first component of the weight is a city’s physical geographical constraints. Cities vary in their endowment of developable land. A city surrounded by water, like San Francisco, has much less developable land than a city surrounded by flat plains, like Houston. Saiz (2010) built an index of developable land for major U.S. cities. His index is constructed in the following manner: a 50 km circle is drawn around the city center. Then, he calculates what percentage of this circle cannot be built upon. Undevelopable land for construction includes oceans, wetlands, lakes, and steep slopes.¹¹ Saiz’s index shows a wide variation of developable land between cities. The most constrained city is Ventura, CA, with only 20 percent developable land, while the least constrained city (McAllen, TX) has over 99 percent developable land. The amount of developable land will be an exogenous component of a city’s housing supply elasticity.¹² Cities with a high percentage of developable land can better respond to demand shocks than cities with low percentages of developable land. Differences in geography should not affect marriage in any plausible way other than through housing. The presence or absence of lakes, oceans, and mountains should not directly induce or dissuade couples from marriage.

The second weight is a population-based weight. A population growth weight is the appropriate weight because permits represent the flow of new housing construction, rather than the existing stock. Current population growth cannot be used as the weight because it will be correlated with contemporaneous demand shocks. Lagged population growth, while uncorrelated with current demand shocks, is problematic because population growth is highly serially correlated. Therefore, I build a predicted population growth weight based on factors that will be orthogonal to short-term demand shocks. Due to the advent of air conditioning, warmer cities began growing faster than colder cities after 1920. Thus, average annual temperature in conjunction with region dummies can be used to predict population growth. I generate a weight called predicted population growth for each of the j cities that is the share of total population growth that city j will account for as predicted by its region and average annual tempera-

¹¹Land with above a 15% slope is classified as steep and difficult to build on.

¹²Saiz (2010) calculates the elasticities and finds constrained cities have lower supply elasticities than less constrained cities.

ture. I then lag the weight by ten years. For example, permits in 1950 will be assigned based on predicted population growth from 1930 to 1940.

The weights are used to construct an instrument (predicted permits) for annual actual permits by the following equation:

$$\hat{P}_{jt} = \left(\sum_j P_{jt} \right) \frac{pgrow_{jt-10} dev_j}{\sum_j pgrow_{jt-10}}$$

The first term is the sum of all permits from j cities in year t . The second term is the predicted population growth weight for city j . The final term (dev_j) is the percent of developable land in city j . Figure 3, a plot of predicted permits per capita against actual permits per capita, shows that the generated predicted permits are highly positively correlated with actual permits for the years 1944 to 1947 (the main years used in the empirical analysis).

5.3 Results

Table 2, the results of estimating (1), shows that building permits per capita from 1947 to 1949 had a significant effect on an individual's probability of marriage in 1950. The OLS estimates suggest that a standard deviation increase in the per capita permits from 1947-1949 increases the probability an individual married in the previous three years by 11 percent for males and 13 percent for females. Columns (2)-(4) report the IV estimates. When lagged population growth is used as the population weight (column 2) in the construction of the instrument, the IV and OLS estimates are statistically indistinguishable. As discussed in 5.2, lagged population growth might not be an appropriate weight; thus column (3) reports the results for when predicted lagged population growth is used as the population weight. These IV estimates are significantly larger than the OLS estimates: a standard deviation increase in building permits per capita raises the probability of marriage over a three-year period by 24 percent for men and 21 percent for women. The difference in magnitude between the OLS estimates (column 1) and the IV estimates (column 3) suggests that measurement error was a larger bias in the estimated effect of building permits than omitted variables. The sample used to estimate equation (1) is a cross section, therefore, I can use the time-invariant weights (geography and lagged predicted population growth) from the construction of the annual instrument as separate instruments. The advantage of using multiple instruments is that the first stage is no longer exactly identified and I can test for the endogeneity of

the instruments with over-identification tests. Column (4) reports the results of using geography interacted with region as one instrument and lagged predicted population growth as a second instrument. The estimated effect of permits on marriage in column (4) is not significantly different from column (3) and the instruments pass the over-identification test, which suggests that the instruments are valid.

Table 3, the results of estimating (2), shows that the estimates of the effect of building permits per capita on marriage probabilities are similar between the cross section and fixed effects specifications. In the fixed effects model, the OLS estimates suggest that a standard deviation increase in permits per capita in the previous two years raises the probability of marriage by 7 percent for men and 8 percent for women over a two-year period. As in the cross section, the difference between the OLS and IV estimates with lagged population growth as the weight are not significantly different. When lagged predicted population growth is used as the weight in the construction of the IV, the estimated effect of permits per capita on marriage outcomes doubles. The estimates in column (3) and (6) suggest that a standard deviation increase in permits raises the probability of marriage by 13 percent for men and 16 percent for women over a two-year period. This effect is quantitatively similar to the 24 and 21 percent effect observed over a three-year period in the cross-sectional results. The similar effect of permits per capita on marriage outcomes between the fixed effects and the cross-section specifications suggest reverse causality was not a significant bias in the estimates from table 2.

After 1950, I perform the empirical analysis with county level data. The county level data has some drawbacks compared to the individual level data. The number of observations in these data is smaller than the census, and the variable of interest (marriage licenses) is measured with more error. People may often file for marriage in a different location than where they reside. Therefore, the marriage license counts may not be a perfect measure of marriage in a given area.¹³ However, measurement error in building permits is less of an issue in the county sample because the permit series sampling method improves after 1954. As in the individual level data, building permits positively influence marriage. Table 4 shows the results of estimating equation 3. At the county level the size of the effect of permits on marriage is larger than at the individual level. The OLS estimates (column 1) suggest that a standard deviation increase in permits

¹³However, if the measurement error is mean independent of the explanatory variables, this issue will not bias the results.

per capita in a given year raised the marriage licenses per capita by 6 percent in the following year. The IV estimates (columns 2 and 3) are larger in magnitude but not statistically different than the OLS estimates, perhaps because there is less measurement error in the permit series.

5.3.1 Entire Period (1926-1949)

I can repeat the estimation of equation (2) from 1928 through 1949 because building permit data are available from the mid-1920s. Like the 1950 census, the 1940 census has both marriage and location information. In the 1940 census there is a question on age of marriage and also a question about an individual's current location and the individual's location five years prior. From these two questions, I construct retrospective marriage histories for persons in this census. I combine the building permit data for a given city with the individual marriage data to estimate the effect of permits on marriage outcomes from 1928 to 1949.¹⁴ Table 5, the results of estimating equation (2) on the whole period (World War II excluded), shows that building permits per capita have the predicted positive effect, but the effect is slightly weaker than it was in the post-war period. The estimates from columns (1) and (2) suggest that a standard deviation increase in permits raises the probability of marriage in a two-year period by 6.3 percent for men and 5 percent for women. The lower effect is likely due to the fact that young people were less apt to own a home prior to the GI bill and thus were less sensitive to permit fluctuations. Furthermore, the variance in permits per capita between cities was lower in the 1930s as the Depression slowed new construction everywhere.

The instrumental variable approach cannot be used prior to 1940. I use ten-year lagged population growth as predicted by temperature in the construction of the instrument. Because air conditioning technology only began to diffuse after 1920, the first decade for which temperature predicts city growth is 1920 to 1930. Therefore, I cannot construct the instrument prior to 1940. The instrument constructed with simple lagged population growth as the population weight is also problematic. From 1930 to 1945 there is very little new building, and many cities lost population. Recall, the instrument is based on two weights: geography and population growth. If cities are contracting, the geography weight will not predict permits because geography will not be a check on

¹⁴I choose 1928 as the cut-off because going further back introduces significant migration bias in the individual data and the permit series is not complete until the late 1920s. However, the results are robust to different cut-off years.

construction when so little is being undertaken. Figure 4, a plot of the correlation between actual and predicted permits by year, shows that the instrument is not valid for the entire period. The instrument is only correlated with actual permits when the economy is expanding in the late 1930s and late 1940s. However, the IV estimates from the postwar period were always equal to or larger than the OLS estimates. Thus, while I cannot use the IV analysis for the full period, it is reasonable to assume that the OLS estimates are a lower bound for the effect of permits per capita on marriage outcomes from 1928 to 1949 and that potential biases are not causing the observed positive relationship between housing supply and marriage.

5.3.2 Effects on Different Demographic Groups

The conceptual framework in Section 3 has clear implications about the effects of housing on different demographic groups. Firstly, housing costs should be more important at younger ages because older men are more likely to have built up savings and would be less sensitive to fluctuations in the housing market. In table 6, panel A, columns 1 and 2 show the results of the estimation of equation (1) on younger men (age 20-27), and columns 3 and 4 show the effect for older men (28-35). The OLS estimate of the effect of permits on marriage is 40 percent smaller for older men, and the IV estimate is 25 percent smaller. While the estimates are not statistically different, they suggest that housing cost matters less for older men.

In the simulation of the model, the difference in marriage rates between low housing cost cities and high housing cost cities is driven by the ability of men from the lower end of the distribution to marry in low housing cost cities. Thus, men from the lower end of the distribution should be more affected by housing costs than men at the higher end of the distribution. Intuitively, a high quality man should be able to obtain housing regardless of which housing cost regime he is in, while a low quality man can obtain housing only when costs are low. One dimension of quality is education. Table 6, Panel B, reports the estimated effect of building permits per capita on men with a high school education versus men with less than a high school education. As predicted by theory, the effect of permits per capita on marriage rates for low education men is around 50 percent higher than the effect for high education men.

The model also predicts that people should marry younger in low housing cost regimes. Table 6, Panel C, shows the effect of building permits on men and

women's age at marriage. Building permits per capita negatively impact the age of marriage for both men and women, although the OLS estimate for women is not statistically significant. The IV estimates suggest a standard deviation increase in permits per capita lowers the age of marriage for men by 4 months and for women by 6 months.

Another interesting group is African-Americans. In the 1950s, African-Americans had to wait for whites to filter out of new housing before they could purchase homes (Boustan and Margo, 2011). Therefore, the effect of new permits should be muted for blacks in this time period. Table 6, Panel D, shows the estimates of equation 1 for whites and blacks. The OLS estimates are 40 percent smaller for blacks versus whites, and the IV estimates suggest that permits per capita had no effect on black marriage outcomes.

5.3.3 Prices

Home prices and building permits have a similar effect on the probability of marriage in 1950. In 2.2, I argue that building permits are a good proxy for the cost of housing in a given city. A more direct proxy for the cost of housing is the median home price within a city. However, while permit data is available annually, home price data is only available for the year 1950. Thus, home prices can only be used in the cross-section specification. Table 7, the results of estimating equation (1) with median home price (in thousands) in 1950 as the variable of interest instead of permits per capita, shows that home prices had a comparable effect to that of permits per capita. The OLS estimates imply that a standard deviation decrease in the median home price raises the probability of marriage by 10 percent for women and 14 percent for men. The IV estimates are about double the magnitude of the OLS estimates: a standard deviation decrease in the median home price raises the probability of marriage by 20 percent for women and 21 percent for men. The fact that a standard deviation change in permits per capita has virtually the same effect on marriage that a standard deviation change in home prices suggests that permits per capita are a valid proxy for the cost of housing.

5.4 Migration

Individual migration is a potential bias in the estimated effect of building permits per capita on marriage outcomes. The census contains the location of persons in 1950 and 1949. In the empirical analysis, I assume that a person was

located in the city reported in 1949 for the years 1946 to 1948. If individuals marry and then move to cities with higher building permits per capita, I will be over-estimating the effect of permits on marriage. Given that high growth cities (the cities with high in-migration) have over double the permits per capita of lower growth cities, the potential bias is indeed positive.¹⁵ However, the majority (66 percent) of persons age 17-27 who moved to a different county from 1949 to 1950 were single. Therefore, it is much more likely that people move and then marry than marry and then move. As a robustness check, I can estimate equation (1) for the years where location of persons is fully known (1949-1950); this sub-sample should have minimal migration bias. Table 8 shows the results of this estimation. The effects of permits per capita on marriage are larger in this sub-sample than they are in table 2, which suggests that migration bias is not driving the main results of the paper. The OLS estimated effect of a standard deviation increase in permits per capita on the probability of marriage is 13 percent for men and 21 percent for women, while the IV estimated effect is 28 percent for men and 33 percent for women.

6 Discussion

In this section, I use my estimates of the effect of housing on marriage to show that growth in the supply of housing can explain nearly 10 percent of the baby boom. Around 750,000 more homes were built between 1947 and 1949 than were built between 1927 and 1929. An increase of this magnitude corresponds to 0.5 more new permits per one hundred people. My estimates suggest that an additional 0.5 permits per hundred persons would raise the marriage probability for women age 18-25 by 4 percentage points and for women age 26-30 by 2 percentage points. Columns (1) and (2) of table 9 report the proportion of women married by age in 1930 and 1950, respectively. Column (3) is my estimate of the proportion of women who would be married in 1930 after the influx of 750,000 new homes. Column (4) is the percentage of the difference between the 1930 marriage rate and 1950 marriage rate accounted for by the housing increase. On average, the new homes can account for about 32 percent of the difference in the proportion married between 1930 and 1950 for women age 20 to 32. According to the demographic exercise I performed in section 2, changes in marriage patterns are responsible for 29 percent of the changes in crude birth

¹⁵I define high growth cities as cities with 1940 to 1950 population growth over the mean and low growth as cities with growth under the mean.

rates from 1930 to 1950. These two estimates taken together suggest that about 9.6 percent of the baby boom can be accounted for by growth in the housing supply after World War II.

Is the above thought experiment realistic? The number of permits from 1927 to 1929 was an equilibrium outcome. If 750,000 new homes were built would they be bought/rented? The key difference between the late 1920s and 1940s was that there was an exogenous innovation in mortgage technology that shifted demand upward. Federal involvement in the home mortgage sector, originally through the Federal Housing Authority in the late 1930s and then perpetuated by the G.I. Bill after World War II, reshaped home loans. The average home loan in 1947, when compared to 1927, had a longer term, lower down payment and lower fixed interest rate. The G.I. Bill gave a segment of the population a power they did not have before: the ability to buy homes. Fetter (2010) estimates that the G.I. Bill raised home ownership rates for 26 year-olds by 70 percent and for 32 year-olds by 32 percent. To accommodate these increases, over a million new homes would have to have been built. Thus, it is feasible that these 750,000 new homes from 1947 to 1949 were built to serve a portion of the population that had not been able to purchase homes prior to the 1944 passing of the G.I. Bill.

7 Conclusion

This paper estimates the effect of housing supply on marriage outcomes during the U.S. baby boom. Specifically, I use permits per capita as a proxy for housing supply and find that a standard deviation increase in permits per capita from 1947 to 1949 increased the probability of marriage by 21-24 percent in that three-year period. My estimates imply that increases in housing supply can explain 32 percent of the difference in marriage rates between 1930 and 1950. The remaining difference can likely be attributed to the strong postwar U.S. economy. In an earlier paper on marriage and the Great Depression, I estimated that a standard deviation increase in GDP raised the probability of marriage by 24 percent over a three-year period (Hill, 2011). Thus, GDP growth and increased housing supply can explain much of the remarkable rise in marriage rates between 1930 and 1950.

I find that growth in the supply of housing can account for 10 percent of the rise in birth rates between 1930 and 1950 through the channel of increased

marriage rates. This result may explain why the baby boom experienced by the U.S. was larger than similar postwar baby booms in other western countries. Most developed countries experienced a baby boom after World War II, but the baby boom was largest in the U.S., Canada, New Zealand and Australia. The urban areas in these four countries had been spared the destruction experienced by European cities and began the post-war period with a largely intact housing stock. Returning soldiers and rising GDPs can probably explain the majority of the baby boom in developed countries. However, increased housing supply may explain why the baby boom was larger in countries with open land and less affected housing stock.

My estimates of the effect of increased housing supply on the baby boom are most likely an underestimate because the growth of housing supply could also account for increased fertility within marriage. For instance, areas with more housing typically have larger dwellings which might enable couples to have more children. Declines in maternal mortality and rising income are among the factors that could explain the additional portion of the baby boom unaccounted for by housing.¹⁶ In future work I will examine the Easterlin hypothesis that cohorts raised during the hardship of the Great Depression preferred to spend their disposable income on children rather than on consumption goods.

¹⁶See Albanesi and Olivetti (2010) and Easterlin (1962).

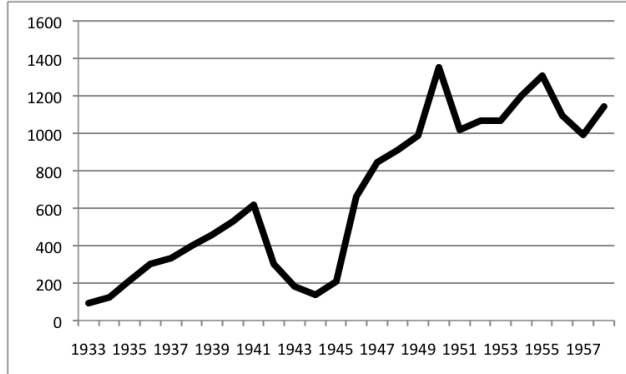
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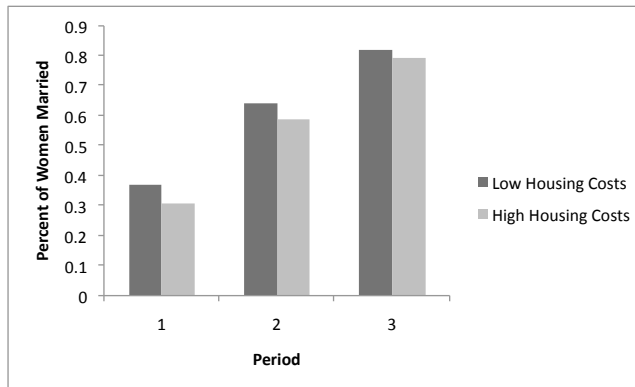
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Figure 1 - Building Permits (1933-1958)



source: Bureau of Labor Statistics

Figure 2 - Housing Costs and Marriage Rates



Notes: the distribution of mates is assumed to be $\sim N(0, 1)$ and $\beta = .9$, $c = .1$ (low cost) and $c = .9$ (high cost)

Figure 3 - Actual vs. Predicted Permits (1944-1947)

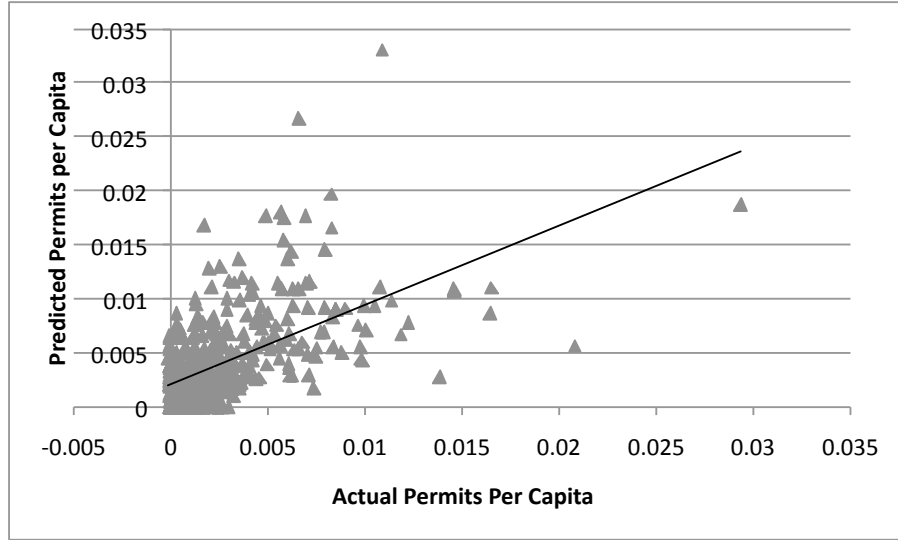
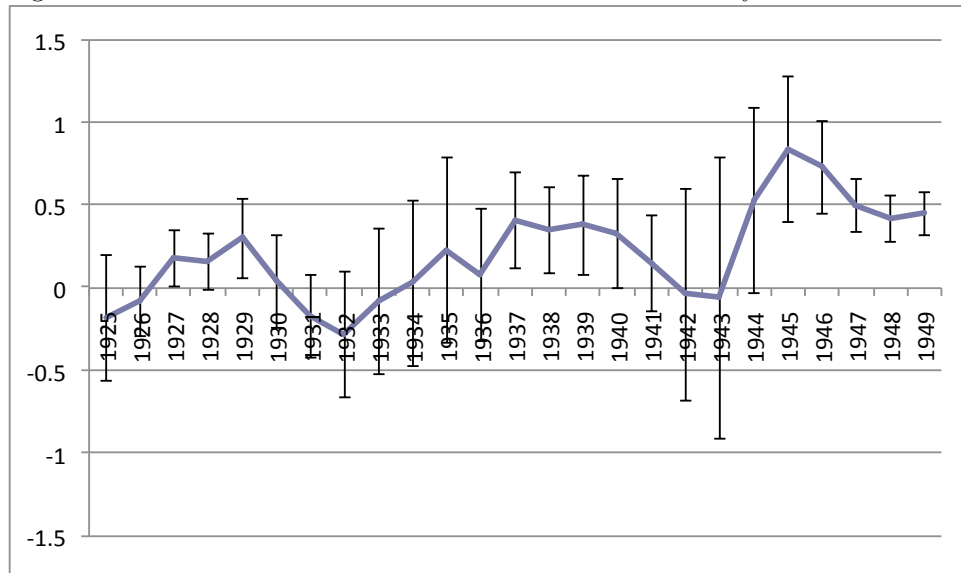


Figure 4 - Correlation between Predicted and Actual Permits by Year



notes: error bars are a 95% confidence interval

Table I - Correlation between Permits and Prices (1950)

Dep. Variable = log Median Home Price in 1950	
log Blding Permits per hundred 1944-1949	-0.092** (.035)
Median Income 1950	0.0003*** (.00007)
Employment Rate	0.649 (.743)
r-squared	0.26
obs	82

Notes: *** significant at the 99% level. ** significant at the 95% level. * significant at the 90% level. Other controls included are housing stock per capita (1940) and average rooms per dwelling.

Table II -Effect of Building Permits on Marriage (1950 cross section)

Panel A: Females				
	Dependent: =1 if married			
	(1)	(2)	(3)	(4)
	OLS	IV(lag)	IV(plag)	IV(2 instr)
Building Permits per hundred (47-49)	0.048*** (.013)	0.058** (.024)	0.093*** (.023)	0.077*** (.015)
Prcnt Change in Retail Sales (40-50)	0.248*** (.051)	0.256*** (.050)	0.284*** (.055)	0.270*** (.053)
Prcnt Change in Employment (40-50)	-0.176 (.346)	-0.040 (.463)	0.459 (.442)	0.224 (.390)
Dummy=1 if in Labor Force	-0.343*** (.017)	-0.343*** (.017)	-0.340*** (.017)	-0.341*** (.017)
r-squared	0.25			
obs	5247	5247	5247	5247
Mean Dep. Var	0.45	0.45	0.45	0.45
Weak Instr. F Test		20.7	10.9	12
p(Over-Id Test)				0.688
Sample	Age 18-25, single/ married in last 3 yrs			
Panel B: Males				
	Dependent: =1 if married			
	(1)	(2)	(3)	(4)
	OLS	IV(lag)	IV(plag)	IV(2 instr)
Building Permits per hundred (47-49)	0.040*** (.014)	0.039** (.018)	0.092*** (.025)	0.072*** (.016)
Prcnt Change in Retail Sales (40-50)	0.219*** (.065)	0.218*** (.065)	0.257*** (.071)	0.242*** (.067)
Prcnt Change in Employment (40-50)	-0.195 (.359)	-0.211 (.441)	0.530 (.581)	0.249 (.382)
Dummy=1 if Veteran	0.136*** (.016)	0.136*** (.016)	0.137*** (.016)	0.137*** (.016)
Dummy=1 if in Labor Force	0.226*** (.019)	0.226*** (.019)	0.227*** (.019)	0.227*** (.019)
r-squared	0.15			
obs	5027	5027	5027	5027
Mean Dep. Var	0.41	0.41	0.41	0.41
Weak Instr. F Test		22.9	11.3	10.5
p(Over-Id Test)				0.363
Sample	Age 20-27, single/ married in last 3 yrs			

Notes: *** significant at the 99% level. ** significant at the 95% level. * significant at the 90% level. Table displays results of estimating equation (1) in the text. Non-whites are excluded. Other controls included are housing stock per capita (1940), sex ratio, education dummy, and foreign-born dummy. Standard errors are clustered by city.

Table III - Effect of Building Permits on Marriage 1946-1949 (Fixed Effects)

	Dep. Variable=1 if Married					
	Male			Female		
	(1) OLS	(2) IV(lag)	(3) IV(plag)	(4) OLS	(5) IV(lag)	(6) IV(plag)
Building Permits per hundred	0.049*** (.013)	0.069*** (.018)	0.090*** (.026)	0.062*** (.017)	0.053** (.022)	0.119*** (.031)
r-squared	0.15			0.24		
obs	10395	10395	10395	10514	10514	10514
Mean Dep. Var	0.4	0.4	0.4	0.42	0.42	0.42
F-stat		20.7	32.2		16.5	28.1

Notes: *** significant at the 99% level. ** significant at the 95% level. * significant at the 90% level. Table displays results of estimating equation (2) in the text. Non-whites are excluded. Other controls included are a labor force dummy, an education dummy, and a foreign-born dummy. Standard errors are clustered by city and period.

Table IV - Effect of Building Permits on Marriage Licenses 1954-1962

	Dep. Var= Marriage Licenses per Capita		
	OLS	IV(lag)	IV(plag)
Building Permits per Capita prev. year	0.101*** (.022)	0.156** (.068)	0.189*** (.073)
F-stat		26.5	22.2
Mean Dep. Var		0.0076	0.0076
N		528	528

Notes: *** significant at the 99% level. ** significant at the 95% level. * significant at the 90% level. Table displays results of estimating equation (4) in the text. Standard errors are clustered by city.

Table V - Effect of Building Permits on Probability of Marriage 1928-1949

	Dep. Var = 1 if Married in period	
	Male	Female
	(1)	(2)
Building Permits per hundred people	0.017*** (.006)	0.016** (.007)
r-squared	0.1	0.16
obs	46748	49312
Mean Dep. Var	0.22	0.25

Notes: *** significant at the 99% level. ** significant at the 95% level. * significant at the 90% level. Table displays results of estimating equation (2) in the text. Non-whites are excluded. Men age 20-27 included, women 18-25 included. Other controls included are a labor force dummy, an education dummy, and foreign-born dummy. Standard errors are clustered by city and period.

Table VI - Effect of Building Permits (by Group)

Panel A: Young vs. Old Men				
	Dep Var. = 1 if Married in last 3 yrs.			
	Men Age 20-27		Men Age 28-35	
	(1)	(2)	(3)	(4)
	OLS	IV(plag)	OLS	IV(plag)
Bldg Permits per hundred (47-49)	0.040*** (.014)	0.092*** (.025)	0.024* (.012)	0.069*** (.021)
Percent Change in Retail Sales (40-50)	0.219*** (.065)	0.257*** (.071)	0.177*** (.051)	0.202*** (.048)
r-squared	0.15		0.05	
F-stat	11.3		10.1	
obs	5027		2727	
Panel B: Low Ed vs. High Ed Males				
	Dep Var. = 1 if Married in last 3 yrs.			
	H.S. Ed or Higher		Less than H.S. Ed	
	(1)	(2)	(3)	(4)
	OLS	IV(plag)	OLS	IV(plag)
Bldg Permits per hundred (47-49)	0.038*** (.011)	0.075*** (.021)	0.055** (.024)	0.117*** (.039)
Percent Change in Retail Sales (40-50)	0.215*** (.058)	0.250*** (.069)	0.213** (.099)	0.272*** (.102)
r-squared	0.16		0.11	
F-stat	11.9		13	
obs	3177		1850	
Panel C: Age of Marriage				
	Dep Var. = Age of Marriage			
	Men, married in last 3 years		Women, married in last 3 years	
	(1)	(2)	(3)	(4)
	OLS	IV(plag)	OLS	IV(plag)
Bldg Permits per hundred (47-49)	-0.163*** (.045)	-0.228** (.116)	-0.133 (.093)	-0.512*** (.166)
Percent Change in Retail Sales (40-50)	-0.976*** (.319)	-1.03*** (.295)	-0.876** (.399)	-1.22*** (.407)
r-squared	0.04		0.06	
F-stat	11.5		10.7	
obs	1778		2031	
Panel D: Black vs. White				
	Dep Var. = 1 if Married in last 3 yrs.			
	White		Black	
	(1)	(2)	(3)	(4)
	OLS	IV(plag)	OLS	IV(plag)
Bldg Permits per hundred (47-49)	0.041*** (.012)	0.080*** (.018)	0.028** (.012)	0.002 (.026)
Percent Change in Retail Sales (40-50)	0.221*** (.046)	0.247*** (.048)	0.204*** (.061)	0.189*** (.061)
r-squared	0.23		0.15	
F-stat	11.4		5.6	
obs	12334		2029	

Notes: *** significant at the 99% level. ** significant at the 95% level. * significant at the 90% level. Table displays results of estimating equation (1) in the text. Other controls included are housing stock per capita (1940) and sex ratio. Age groups are age 20-27 for males and 18-25 for females unless otherwise noted. Standard errors are clustered by city.

Table VII - Effect of Home Prices on Probability of Marriage (1950)

	Dependent: =1 if married			
	Females		Males	
	(1) OLS	(2) IV(plag)	(3) OLS	(4) IV(plag)
Median Home Price	-0.023*** (.0055)	-0.044*** (.013)	-0.028*** (.0061)	-0.042*** (.012)
Dummy=1 if in Labor Force	-0.343*** (.017)	-0.343*** (.017)	0.231*** (.018)	0.234*** (.019)
Dummy=1 if Veteran			0.134*** (.016)	0.133*** (.016)
r-squared	0.25		0.15	
obs	5247	5247	5027	5027
Mean Dep. Var	0.45	0.45	0.41	0.41
Weak Instr. F Test		10.1		10.1

Notes: *** significant at the 99% level. ** significant at the 95% level. * significant at the 90% level. Table displays results of estimating equation (1) in the text. Non-whites are excluded. Other controls included are percent change in retail sales (1940-1950), percent change in employment (1940-1950), housing stock per capita (1940), sex ratio, education dummy, and foreign-born dummy. Standard errors are clustered by city.

Table VIII - Effect of Building Permits on Probability of Marriage (1949-1950)

	Dep. Variable=1 if Married			
	Male		Female	
	(1) OLS	(2) IV(plag)	(3) OLS	(4) IV(plag)
Building Permits per hundred people (47-49)	0.029** (.011)	0.062** (.026)	0.051*** (.015)	0.081*** (.022)
Percent Change in Retail Sales (40-50)	0.164*** (.044)	0.188*** (.053)	0.214*** (.044)	0.237*** (.048)
Percent Change in Employment (40-50)	0.045 (.262)	0.499 (.492)	0.020 (.379)	0.424 (.406)
Dummy=1 if Veteran	0.076*** (.016)	0.077*** (.017)		
Dummy=1 if in Labor Force	0.148*** (.014)	0.150*** (.014)	-0.250*** (.017)	-0.247*** (.018)
r-squared	0.06		0.14	
obs	3878	3878	3930	3930
Mean Dep. Var	0.23	0.23	0.26	0.26
F-stat		10.1		10.7
Sample	Age 20-27, single or married in last 1 yrs Age 18-25, single/ married in last 1 yrs			

Notes: *** significant at the 99% level. ** significant at the 95% level. * significant at the 90% level. Table displays results of estimating equation (1) in the text. Non-whites are excluded. Other controls included are housing stock per capita (1940), education dummy, sex ratio and foreign-born dummy. Standard errors are clustered by city.

Table IX - Proportion of Females Married by Age (30, 50, 30_predict)

	(1)	(2)	(3)	(4)
Age	1930	1950	1930_hat	Percent of difference
20	39.34	46.95	43.34	0.52562418
21	46.59	58.49	50.59	0.33613445
22	54.8	66.68	58.8	0.33670034
23	62.82	74.18	66.82	0.35211268
24	66.98	79.16	70.98	0.32840722
25	72.14	83.5	76.14	0.35211268
26	75.63	85.74	77.63	0.19782394
27	79.26	87.75	81.26	0.23557126
28	81.19	88.85	83.19	0.26109661
29	84.16	91.22	86.16	0.28328612
30	83.21	90.77	85.21	0.26455026
31	87.25	92.89	89.25	0.35460993
32	86.34	92.7	88.34	0.31446541

Notes: Column (3) is the predicted proportion of females married in 1930 if permits were at their 1947-1949 levels. Column (4) is the percent of the difference between (1) and (2) that can be accounted for by the increase in permits.