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Fertility Timing of Unmarried and Married Mothers: Evidence on Variation Across U.S. Cities from the Fragile Families and Child Wellbeing Study

Marah A. Curtis and
Boston University, 246 Bay State Road, Boston, MA 02215, USA mcurtis@bu.edu

Jane Waldfogel
Columbia University, 1255 Amsterdam Avenue, New York, NY 10027, USA

Abstract

In this paper, we examine the determinants of fertility timing of unmarried and married mothers using a rich new birth cohort study, the Fragile Families and Child Wellbeing Study, drawn from 20 medium and large U.S. cities. We find considerable variation in the time to next birth among comparable mothers who live in different cities. Some of this variation is explained by variation in labor markets, housing costs and availability, and welfare policies. City variation is particularly important for unmarried women who already have two or more children, whose fertility is more sensitive to these contextual variables than is the fertility of married women, or unmarried women with just one child.

Keywords

Fertility timing; U.S. city variation in fertility; Policies and fertility

Introduction

Unmarried mothers, their fertility, and family formation patterns are frequently the focus of public policy research and debate. Mothers who are unmarried at the birth of a child are of particular interest to policymakers due to concern about their greater likelihood of poverty and potential reliance on public benefits, as well as concern about the long-run effects on the children (McLanahan and Sandefur 1994; Powell and Parcel 1997; Committee on Ways and Means 2004; Curtis 2007). Even given this interest, however, very little research has looked comprehensively at the impact of policies on fertility decisions and timing among this group of mothers. Information about the relationships between public policies and unmarried mothers' fertility may yield insights into how these fertility decisions are made and what impact policies have on these decisions.

In this paper, we examine the determinants of variation in fertility timing using a new and diverse birth cohort study, the Fragile Families and Child Wellbeing Study (FFCW). The FFCW sample is drawn from 20 medium and large U.S. cities selected to provide variation in labor markets, housing markets, and welfare regimes. The FFCW provides longitudinal information about a birth cohort of 3,712 children born to unmarried parents as well as a comparison group of 1,186 children born to married parents, in 75 hospitals in twenty U.S. cities with populations of 200,000 or more and is representative of births in large cities. Sample

Correspondence to: Marah A. Curtis.

members are followed longitudinally, with the latest wave of data available from approximately 3 years post-birth.

Taking advantage of the rich longitudinal data from the birth cohort study and the extensive variation across cities in labor markets, housing costs and availability, and welfare policies, we analyze the timing to next birth and the influence of these contextual factors on that timing among this large and diverse sample of mothers. We find considerable variation in the time to next birth across cities, some of which is explained by the variation in labor markets, housing markets, and welfare policies, particularly for unmarried women who already have two or more children. Our results indicate that this group's fertility is more sensitive to these contextual variables than is the fertility of married women, or unmarried women with just one child.

Background

Economists, sociologists, and demographers have long been interested in the influence of labor markets, housing costs, and welfare policies on fertility. A large literature dating back to work by Butz and Ward (1979) has examined the influence of labor markets on women's fertility decisions (see review in Hotz et al. 1997). There have also been numerous studies of the effects of welfare on family formation outcomes, including fertility (Fairlie and London 1997; see Moffit 1998 for a comprehensive review). The effects of housing costs have been less studied. We were able to find only one prior study of the effects of housing costs on fertility, a study by Curry and Scriven (1978) who found that apartment dwellers have lower fertility in tight housing markets. The omission of housing costs from studies of fertility is problematic as housing costs have been shown to have strong effects on family formation outcomes such as marriage and cohabitation (Winkler 1992; London 2000; Sigle-Rushton and McLanahan 2002; Hughes 2003; Curtis 2007), thus raising the possibility that they might affect fertility as well.

A second, and related, shortcoming of prior research on fertility is that it has tended to focus on just one set of policies (typically, labor market or welfare policies) rather than analyzing all relevant policies concurrently. Studies that control for only one or two sets of factors, rather than a fuller set, may be biased if those factors are correlated with each other, as well as the outcome variable. For example, since cities with higher welfare benefits have higher benefits in part because the cost of living is higher, the interpretation of the welfare benefit variable is problematic if wages are not controlled (Blau et al. 2004). A similar problem arises if housing costs are not included.

Our study expands upon previous research by including detailed information on labor markets, housing costs and availability, and welfare policies, unlike prior studies that have tended to focus on only one or two of these areas at a time. As noted above, housing costs and availability have been largely neglected as an influence on fertility. This omission is particularly worrisome given that the few studies that have looked at housing costs and family formation outcomes have found them to have important effects. Therefore, we include detailed controls for housing costs and availability in all our models, along with detailed controls for labor markets and welfare policies.

Theoretical Framework

In order to estimate associations between labor markets, housing markets, welfare policies, and women's decisions about the timing of subsequent fertility, we begin with the theoretical framework presented by Walker (1995) (see also Bjorklund 2006). In this framework, policies influence fertility by altering the costs of having a child. Three specific types of costs enter into a woman's decision: (1) direct costs (of housing, food, clothing, etc.) associated with raising the child; (2) foregone current earnings associated with time out of the labor market;

and (3) lower future earnings due to the loss of human capital associated with the loss of work experience and job tenure.

Given this framework, higher housing costs would, all else equal, be expected to decrease fertility by raising the direct costs of having a child (assuming that as children are added to a family, this increases the amount of housing the family requires). Conversely, higher welfare benefits, all else equal, should increase fertility by lowering the direct costs. The effects of labor market factors are less straightforward. Better labor markets, as evidenced by lower unemployment rates and higher wages, would be expected to lower the direct costs of having a child by raising family incomes, but also should increase the penalty to having a child, by raising the value of foregone current earnings as well as the reductions in future earnings (assuming that cities with strong labor markets today will have strong labor markets in future).[1] Higher unemployment could also depress fertility if parents respond by investing more in existing children rather than bearing additional children (Becker et al. 1990;Adsera 2005). Thus, the influence of labor market factors on fertility is not clear a priori.

Another implication of this framework is that the costs of having a child, and the effects of contextual factors, may not be constant across women. The number of prior births is likely to be particularly consequential. Given that the modal family size in the U.S. consists of two children, women who already have two or more children may be more sensitive to costs, and policies that affect costs, than women who have fewer or no children. Therefore, in our sample, where all women have just given birth and thus have at least one child, there may be a distinction between women who have just one child (and may want another to complete their family) and women who already have two or more children (and for whom another child may be more discretionary).

A second factor that may alter the costs of children, and in particular the effects of contextual factors, is marital status. There are several reasons why marital status is likely to be consequential. First, married mothers may have access to increased credit to purchase housing conducive to larger families (Leland 2008). Low interest rates and expanded access to credit may both make attractive and increase the demand for owner-occupied housing. Increased demand drives up housing costs suggesting a positive relationship between housing costs and fertility for married women. Unmarried women, however, given single earner status likely will not benefit from increased access to credit to purchase housing and may decrease fertility under higher costs. Thus, the impact of housing costs on married and unmarried women may differ. Second, to the extent that unmarried women bear more of the costs of children than do married women, the fertility decisions of women who are not married may be more sensitive to the unemployment rate and wages. Third, child support enforcement and income-conditioned benefits like Temporary Assistance for Needy Families (TANF) and Food Stamps will tend to be more consequential for unmarried mothers who are more likely to qualify for these benefits. Fourth, the availability of subsidized housing directly affects the stock of low-income housing which will mainly be available to unmarried mothers (although it also, to some extent, affects the overall housing market). This is therefore another instance where we would expect different impacts by marital status.

[1]There is likely to be some asymmetry with regard to gender, if better labor market conditions promote fertility by raising men's ability to support children but also deter fertility by increasing the costs to women of having children. We test for this by including measures of labor market conditions for women as opposed to men although these measures are likely endogenous. In results not included but available upon request, women's labor market measures are not significantly related to fertility. We find these results unconvincing and likely insignificant due to endogeneity. In results including both male and female labor market measures substantial collinearity is introduced while estimates remain insignificant for all measures. We, therefore, use measures of men's unemployment and wages as measures of the overall strength of the labor market, recognizing that the interpretation of their effects will not be straightforward.

Data and Methods

We use data from the Fragile Families and Child Wellbeing Study (FFCW), a unique new national survey that provides longitudinal information about a birth cohort of children born to unmarried parents as well as a comparison group of children born to married parents, in 75 hospitals in 20 U.S. cities with populations of 200,000 or more.[2] We use 19 of these cities.[3] Mothers were interviewed in the hospital shortly after their child's birth and approximately 1 year and 3 years later.[4] Baseline interviews took place for 13 of the cities in 2000, five of the cities in 1999 and two of the cities in 1998.[5]

Our focus is on whether a woman has had a next birth and the timing to that birth. We track subsequent births using data from the household roster that interviewers complete in phone interviews conducted approximately one and three years post-birth. FFCW began with a sample of 4,898 births (3,713 non-marital and 1,185 marital). Of these, 4,231 women (3,181 who had non-marital births and 1,050 who had marital births) remained in the study at the 3-year follow-up. Thus, the attrition rate was 14% overall (14% for non-marital births and 11% for marital).[6]

We begin by categorizing the women in the sample by whether they had any subsequent birth between the time of the baseline birth and the 3-year follow-up survey.[7] We then estimate Cox proportional hazard regression models of the risk of a subsequent birth.[8] Cox proportional hazards models have been widely used in studies of fertility timing (see, for example, Pong 1994;Acs 1996;Finer and Zabin 1998) and are appropriate for this analysis because no assumptions are made about the underlying distribution of event times, the shape or the age/time profile of the baseline hazard. Because our period of observation ends at the 3-year follow-up, mothers who did not have a subsequent birth by that time are right censored. Mothers who had more than one subsequent birth between the baseline birth and the 3-year follow-up (this applies to 121 women, or 0.03% of our sample) are used only once (that is, we analyze the timing to their first subsequent birth and do not analyze the timing to a second subsequent birth, if present).

Because, as discussed earlier, the timing to a next birth, and the determinants of that timing, may vary both by the birth order of the baseline birth, as well as the mother's marital status at the time of that birth, we estimate models separately by both baseline birth order and marital status. Thus, we estimate models for four distinct groups: (1) non-marital first birth at baseline; (2) marital first birth at baseline; (3) non-marital higher-order birth at baseline; and (4) marital higher order birth at baseline. Each model includes controls for labor markets, housing costs, housing subsidy availability, and welfare policies. A strength of the FFCW dataset is that the sample cities were specifically selected to provide a range of labor markets, housing markets,

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[2]Austin, TX; Baltimore, MD; Boston, MA; Chicago, IL; Corpus Christi, TX; Indianapolis, IN; Jacksonville, FL; Nashville, TN; New York, NY; Norfolk, VA; Philadelphia, PA; Pittsburgh, PA; Richmond, VA; San Antonio, TX; San Jose, CA; Toledo, OH; Detroit, MI; Milwaukee, WI; Newark, NJ and Oakland, CA.

[3]We exclude Jacksonville (N = 60) because there were no marital births in that city.

[4]A further follow-up interview was held when the child was about 5 years old but those data are not yet publicly available for analysis.

[5]Corpus Christi, Indianapolis, Milwaukee, New York, San Jose, Boston, Nashville, Chicago, Toledo, San Antonio, Pittsburgh and Norfolk baseline interviews occurred in 2000; Baltimore, Detroit, Newark, Philadelphia and Richmond in 1999 and Oakland and Austin in 1998.

[6]Cases lost to attrition were significantly more likely to be non-marital, foreign born, not working, Hispanic and less educated while there were no differences in terms of age and number of children at baseline.

[7]The FFCW data do not allow us to track pregnancies that did not result in a birth.

[8]The Cox proportional hazard model assumes the effects of each covariate on the hazard rate are the same at any event time. Following methods noted in Grambsch and Therneau (1994), we test this assumption using the "estat phtest" function in STATA and find that only a few covariates (monthly church attendance and having a male focal child) have different associations with the risk of fertility at different points in time. We also tested alternate models interacting time periods with wage (computed on the risk set corresponding to substantively defensible event times) and found the results inconsistent suggesting that the preferred model should only include wage without interactions.

and welfare regimes, so that researchers could estimate the effects of those factors on family formation and other outcomes.

We measure the strength of the labor market with two controls, the male unemployment rate and the natural log of the mean male wage, both derived from data on prime age men (age 25–54) in the March Current Population Survey for that state in the baseline interview year. The male unemployment rate is the number of males aged 25–54 who reported being either unemployed and looking for work or unemployed on layoff divided by the number of males of the same age who reported either working or being with a job but not at work. The log of male wages is the mean hourly wage of males 25–54. As discussed earlier, the effect of these measures is unclear a priori, as a higher unemployment rate and lower wages would be expected to increase fertility by lowering the costs associated with foregone earnings and lower human capital, but would also decrease fertility by raising the costs of children relative to families' incomes.

We measure housing with three variables. The first is a control for housing costs, which we measure with the log of the house price index from Malpezzi, Chun and Green's 1990 single owner occupied housing indexes (Malpezzi et al. 1998) adjusted to 2000 figures using the Office of Federal Housing Enterprise Oversight house price indexes.[9] The second is a measure of the availability of subsidized housing, constructed from HUD's "Picture of Subsidized Households 1998" (U.S. Department of Housing and Urban Development 2005) and "The Low Income Housing Tax Credit Database" (U.S. Department of Housing and Urban Development 2006). This measure includes project based assistance in the form of public housing and tenant-based assistance in the form of certificates and Section 8 vouchers as well as the total number of low-income housing tax credits units. This measure of the number of subsidized units is tabulated by MSA, matched to the Fragile Families cities, and appended to the microdata. We transform this measure to reflect the availability of subsidized units, by dividing the number of subsidized units by the population in a given city, using population data extracted from the U.S. Census Bureau, County and City Data Book. Theory would predict that greater availability of subsidized housing and lower house prices would increase fertility by lowering the costs of having children. However, there may also be interactive effects of these two factors, since greater availability of subsidized housing should have a larger impact on fertility when housing prices are higher. Therefore, in addition to the two main effects for housing cost and availability, we also include an interaction term, which is the product of the two. This is our third housing variable.

We measure welfare policies with two variables: a control for the log of the maximum combined welfare and food stamp grant available to a family of four in the state and year through the Temporary Assistance to Needy Families (TANF) and Food Stamps (FS) programs, using data from the Welfare Benefits Data Base (Moffitt 2005); and a control for the log of child support expenditures per single-mother family in the state and year, using data on child support enforcement expenditures from the Office of Child Support Enforcement and data on the number of single mother families from the Census. Higher welfare and food stamp benefits should raise fertility by lowering the direct costs of children. The effects of child support expenditures are harder to predict. On the one hand, tougher enforcement should raise women's incentive to have children by increasing the likelihood that they will get support and the amount of support collected even if the father is absent, but on the other hand, tougher enforcement should raise men's incentives to avoid having children, because it increases the costs of children for those who are not married or coresident. The net effects will depend on

[9]The Office of Federal Housing Enterprise Oversight (OFHEO) house price index is a weighted repeat sales index designed to measure changes in single-family home values in the U.S. The indexes are adjusted by normalizing the OFHEO housing price index to 100 in the first quarter of 1990 and then multiplying the 1990 value by the 2000 index and dividing by 100.

the relative strength of these contrary effects for men and women and how they balance each other out (Nixon 1997; Willis 1999).

We also include a full set of controls for demographic characteristics that theory and prior research have indicated affect fertility timing. These characteristics, all measured at the time of the baseline birth, include: the mother's age (dummy variables for under age 20, age 25–29, or age 30 or more, with age 20–24 the reference category); mother's race/ethnicity (dummy variables for White, non-Hispanic, Hispanic, or Other, with the reference category Black, non-Hispanic); mother's nativity (dummy variable for foreign-born); mother's education (dummy variables for less than high school or high school, with more than high school the reference category); mother's employment (dummy variable for mother worked in the past year); mother's religious observance (dummy variables for attending services weekly, monthly, or yearly, with hardly or never attend the reference category); child gender (dummy variable for whether child born at baseline is male); and child low birth weight (dummy variable for whether the child born at baseline weighed 5.5 pounds or less at birth).

As noted above, we estimate separate models by number of children at baseline as well as marital status at baseline. About 38% of our sample had just one child at baseline. For the remaining 62% who had two or more children at baseline, we also include in our models a control for the number of children at baseline (dummy variable for having two children, with three or more children the reference group). Approximately 24% of our sample was married at baseline. The 76% who were not married at baseline includes women who are cohabiting with the child's father, in a visiting relationship with the child's father, or not currently seeing the child's father. We control for baseline co-residence status among the unmarried (dummy variables for cohabiting and visiting, with no relationship the reference group). We do not control for subsequent marriage or relationship status because this is likely to be simultaneously determined with subsequent fertility.

Results

Descriptive Statistics

About a quarter of mothers have had a subsequent birth by the time of the 3-year follow-up interview with a mean time of 29 months to birth (Table 1). More than three quarters of the sample are unmarried, 30% having their first birth at baseline and 46% having a second or higher order birth at baseline. Nearly a quarter of the sample are married, 9% having their first birth at baseline and 16% having a second or higher order birth at baseline. The share with a subsequent birth is nearly identical for unmarried mothers (26%) and married mothers (25%) with a mean time of 29 months to birth for unmarried mothers and a mean of 30 months to birth for married mothers. The share of married mothers with a first birth at baseline who move on to a subsequent birth is 38% compared to only 18% for married mothers who already had two or more children at baseline. The share of unmarried mothers with a first birth at baseline who have a subsequent birth as well as those with a second order or higher birth at baseline is nearly identical at 25% and 26%.

As shown in Table 2, the characteristics of those having a subsequent birth do vary by marital status, suggesting that the determinants of a subsequent birth may not be the same for unmarried and married women. The age of the woman seems to matter for subsequent fertility, regardless of whether she is married or unmarried, with younger women more likely to have a subsequent birth and older women less likely. But other determinants differ. Among married women, the number of children at baseline differs significantly between women who go on to have a subsequent birth and those who do not, but this is not the case for unmarried women, for whom we see significant differences in subsequent fertility by maternal education and nativity, neither of which are significant for married women.

Mothers vary not just by whether they have had a subsequent birth but also by the time that elapses before that birth. Kaplan-Meier estimates of the survivor function for time without a birth (Fig. 1) indicate that by 20 months post-baseline, only about 10% of the sample has had a subsequent birth. The share with a birth increases to about 20% by 30 months, and about a third by 40 months.[10] At any point, the slope of the survivor function reflects the magnitude of the hazard. Figure 2 graphs the survivor function for women who had at least two children at baseline, and shows a gap in the subsequent fertility timing of married mothers and unmarried mothers, with unmarried mothers proceeding to a next birth at a much faster rate. At 30 months, for example, over 20% of unmarried mothers of two or more children have had a next birth, as compared to about 15% of married mothers with two or more children. This gap widens over time, so that by 40 months, over 30% of the unmarried mothers have had a next birth, as compared to just over 20% of the married mothers. (Estimates that consider separately women who had just two children at baseline as opposed to those who had more than two children at baseline yield a similar pattern of results; therefore, we combine the two groups.)

As noted earlier, our sample is drawn from 19 U.S. cities that vary a good deal in terms of their labor markets, housing costs, and welfare policies (Table 3). Among the contextual factors, the availability of subsidized housing appears to vary most markedly across cities, ranging from a low of about 0.015 units per person (in Austin) to a high of 0.151 (in Pittsburgh). This variation reflects the historically large role that state and localities have played in the provision of public housing and the different decisions large cities have made about how much to invest in public housing (Olsen 2002;O'Flaherty 2005).

Results from Proportional Hazards Regression Models

To explore the cross-city variation in the risk of a subsequent fertility, we estimate a series of Cox proportional hazards models, including controls for the demographic characteristics described above along with a set of city fixed effects. We estimate models separately by baseline marital status and baseline number of children. Thus, as discussed earlier, we examine four groups: unmarried women with one child, unmarried women with two or more children, married women with one child, and married women with two or more children.

The results for women who had two or more children at baseline, shown in Table 4, confirm the suggestion in the raw data that the determinants of subsequent fertility do vary by baseline marital status. In particular, the city fixed effects exert a stronger influence for the non-marital group than the marital group. For the non-marital group, 8 of the 18 city fixed effects are statistically significant (at $p < 0.01$ or $p < 0.05$) or marginally significant ($p < 0.10$), versus only 1 of 18 for the marital group. Results from comparable models for women who had just one child at baseline, not shown but available upon request, suggest that city effects play a much less important role for these women. Among women who had just one birth at baseline, only five city fixed effects were significant for the non-marital group and two for the marital group. These results suggest that to the extent fertility timing varies by cities in this sample, that variation is mainly driven by variation in the risk of births to unmarried women who already have two or more children.

To explore the role of labor markets, housing costs, and welfare policies in explaining this variation, we estimate another set of models for unmarried women who had two or more births at baseline, adding controls for these contextual factors to our models in place of the city fixed effects.[11] Since mothers are clustered in cities, all models use robust standard errors.

[10]Tied events are handled using the Breslow (1974) approximation of the exact marginal. The risk pools for the second and following failure events within a group of tied failures are not adjusted for prior failures.

The results, shown in Table 5, indicate that a number of these contextual variables have significant effects on subsequent fertility timing. Higher male wages are associated with a significantly lower risk of subsequent birth, consistent with an interpretation that they reflect a better labor market and thus higher costs for women who have children in terms of foregone earnings. Higher male unemployment is also associated with a significantly lower risk of subsequent birth, consistent with an interpretation that it may reduce men's ability to support children (and thus raise the costs of children relative to family incomes). Increased expenditures on child support enforcement are associated with a significantly higher risk of subsequent fertility, consistent with an interpretation that access to child support payments increases a mother's ability to support children on her own and thus her willingness to have additional children. However, higher welfare and Food Stamp benefits have no significant effect on fertility, a result also found by Fairlie and London (1997).[12] In order to ensure that the contextual policy variables are not actually reflecting hospital level variation we estimated additional models including birth hospital fixed effects. In results not shown but available on request, controls for birth hospital fixed effects are not significant as a group nor does their inclusion alter the pattern of the policy effects.[13]

As discussed earlier, we include three controls for housing: a measure of the availability of subsidized housing, a measure of housing costs, and an interaction term (housing costs times availability of subsidized units), since the importance of the supply of subsidized units probably varies by the cost of housing. All policy variables are mean centered. At the mean, main effects for the availability of subsidized housing and housing costs are both insignificant. Costs and subsidizes cannot be understood in isolation but only interactively. The interaction effect of housing costs and availability of subsidized housing is positive and significant, suggesting that fertility is higher when more subsidized housing is available and housing costs are higher than the mean. We conduct F-tests for the joint significance of our policy variables (unemployment rate, wages, subsidized housing, housing costs, housing costs*subsidized housing, TANF/FS grant and expenditures on child support enforcement) and find that as a group they are significant ($p = 0.04$).

Table 6 shows results of similar models estimated for unmarried women who had a first birth at baseline. Similar to results for unmarried mothers with a higher order birth, main effects of subsidized housing and housing costs are insignificant while the interaction effect of housing costs and availability of subsidized housing is positive and significant. However, in contrast to the results for unmarried mothers with two or more children, unemployment rates, male wages and expenditures on child support enforcement are not significant factors for this group. For these mothers, housing costs and the availability of subsidies matter when housing costs are high and more subsidies are available. We conduct F-tests for the joint significance of our policy variables and find that as a group they are significant ($p = 0.00$). In results not shown but available on request, we repeated these models for the two groups of married women: those with one child at baseline, and those with two or more children at baseline. Our earlier models for these groups had found mostly insignificant city fixed effects, suggesting that varying city

[11]With only one cohort observed per city, we cannot control for city fixed effects as well as measures of labor markets, housing costs, and welfare policies. Although in principle these variables could change over the 3-year follow-up period, in practice the change over time is minimal. So essentially, we have to treat these as time-invariant factors.

[12]We use baseline marital status in our analyses, however, mothers may marry sometime between baseline and a subsequent birth. It is possible that the effect of policies may be different when marital status at the next birth is used rather than at baseline. In results not included, but available upon request, our main findings for unmarried mothers with a subsequent higher order birth are quite similar to those reported in Table 5. The risk of a subsequent non-marital birth is reduced when unemployment rates, wages and housing costs increase while the availability of subsidized housing and increased expenditures on child support enforcement are associated with an increased risk of birth. Results, therefore, are not particularly sensitive to our choice to define marital status at baseline rather than at the subsequent birth.

[13]The policy effects are not sensitive to a number of different fixed effects specifications omitting different hospitals or subsets of hospitals.

policies are not likely to be particularly important to these groups. The regression results with the policy variables bear this out.

We also estimated alternative models in which we controlled in more detail for specific welfare policies that might be particularly consequential for fertility (policies such as the family cap, which limits welfare benefits when an additional child is born). However, the additional policies rarely had significant effects, and controlling for them did not alter the overall pattern of results.

Simulation Results

So how consequential are contextual differences across cities in explaining the variation in subsequent fertility among the women whose fertility varies most sharply—the women who are unmarried at baseline and already have two or more children? To shed light on this, we conducted a simulation where we held the characteristics of women constant and then altered the policy, labor market, and housing environments they faced. Taking women from Milwaukee as the base case (because they had one of the highest rates of subsequent fertility), we estimated what their survivor function would be if instead of facing the Milwaukee context, they faced the context of other sample cities.

We illustrate the results in Fig. 3, showing how the survivor function for unmarried women with two or more children at baseline changes as we move the environment from Milwaukee to Oakland, then Chicago, and finally New York. Moving from Milwaukee to Oakland increases the share of women without a birth at 30 months from 75% to 83%; moving to Chicago or New York further increases that to 87%. This means that the share with a subsequent birth is cut in half, from 25% to 13%.

To get a sense of which specific contextual factors may be most important, we perform a series of simulations asking what happens to the survivor function at 30 months for unmarried women with two or more children at baseline if each factor, in turn, is set to the New York level while all other factors are maintained at Milwaukee levels. Unemployment rates, wages, the availability of subsidized housing, welfare and food stamp benefits, and child support enforcement levels are quite similar in Milwaukee and New York. Housing costs, however, are markedly higher in New York, thus this city provides an interesting comparison case.

As we saw in Fig. 3, moving unmarried mothers with two or more children at baseline from Milwaukee to New York increases the share of these mothers without a subsequent birth by 12% points. The simulation results (Table 7) give a sense of the relative importance of factors driving the differences across the cities. Taking each policy individually, housing costs and expenditures on child support enforcement are most important, increasing the survivor function about 6% points each, followed by wages and subsidized housing, while unemployment rates and welfare and food stamp benefits are the least important. The interaction term for housing costs and availability of subsidized housing has an important offsetting effect: moving from the Milwaukee to New York value for this variable is associated with a six point *reduction* in the proportion of mothers without a birth at 30 months to 69%.

Conclusions

This paper provides new evidence on the role that labor markets, housing costs and availability, and welfare policies play in the fertility decisions of women in U.S. cities. Taking advantage of rich new data on mothers in 19 medium to large U.S. cities, our major finding is that, even accounting for a number of important personal characteristics associated with fertility, policy and local area conditions make a difference. Moreover, not only do these contextual factors matter, they appear to matter mostly to unmarried mothers who presumably face a higher cost of child rearing than married mothers. We find that unmarried mothers moving to their second birth are sensitive to housing variables, while unmarried mothers moving to a third or higher order birth are affected by unemployment, wages, and child support enforcement, as well as housing variables. Simulation results for this latter group suggest that their fertility is most affected by housing costs and expenditures on child support enforcement followed by wages, subsidized housing, and unemployment rates.

A perhaps surprising finding is that the generosity of welfare benefits does not play an important role for either group of unmarried mothers. This finding suggests that future research on the effects of policies on fertility and family structure should broaden its focus from welfare benefits and instead examine a larger set of contextual factors including other policies (such as child support enforcement) as well as labor markets and housing costs and availability. Our results suggest that factors related to housing are likely to be particularly consequential.

References

Acs G. The impact of welfare on young mother's subsequent childbearing decisions. The Journal of Human Resources 1996;31(4):898–915. doi:10.2307/146151.

Adsera A. Where are the babies? Labor market conditions and fertility in Europe. 2005:1576. Institute for the Study of Labor (IZA). Discussion paper No.

Becker GS, Murphy KM, Tamura RF. yHuman capital, fertility, and economic growth. The Journal of Political Economy 1990;98:S12–S37. doi:10.1086/261723.

Bjorklund A. Does family policy affect fertility? Lessons from Sweden. Journal of Population Economics 2006;19:3–24. doi:10.1007/s00148–005–0024–0.

Blau FD, Kahn LM, Waldfogel J. The impact of welfare benefits on single motherhood and headship of young women: Evidence from the census. The Journal of Human Resources 2004;39(2):382–404. doi: 10.2307/3559019.

Breslow N. Covariance analysis of censored survival data. Biometrika 1974;30:89–99.

Butz, W.; Ward, MP. Labor markets and fertility: A demographically disaggregate model of US postwar experience. Santa Monica: RAND; 1979.

Committee on Ways and Means. 2004 Green book. Washington, DC: U.S. Government Printing Office; 2004.

Curry JP, Scriven GD. The relationship between apartment living and fertility for blacks, Mexican-Americans, and other Americans in Racine, Wisconsin. Demography 1978;15(4):477–485. doi: 10.2307/2061200. [PubMed: 738474]

Curtis MA. Subsidized housing, housing prices and the living arrangements of unmarried mothers. Housing Policy Debate 2007;18(1):145–170.

Fairlie R, London RA. The effect of incremental benefit levels on births to AFDC recipients. Journal of Policy Analysis and Management 1997;16(4):575–597. doi:10.1002/(SICI)1520–6688 (199723) 16:4<575::AID-PAM4>3.0.CO;2–D.

Finer LB, Zabin LS. Does the timing of the first family planning visit still matter? Family Planning Perspectives 1998;30(1):30–42. doi:10.2307/2991523. [PubMed: 9494813]

Grambsch PM, Therneau TM. Proportional hazards tests and diagnostics based on weighted residuals. Biometrika 1994;81:515–526. doi:10.1093/biomet/81.3.515.

Hotz, JV.; Klerman, JA.; Willis, RJ. The economics of fertility in developed countries. In: Rosenzweig, MR.; Stark, O., editors. Handbook of population and family economics. Vol. Vol. 1. Amsterdam: Elsevier; 1997.

Hughes ME. Home economics: Metropolitan labor and housing markets and domestic arrangements in young adulthood. Social Forces 2003;81(4):1399–1429. doi:10.1353/sof.2003.0059.

Leland J. From the housing market to the maternity ward. New York Times. 2008 February 1;

London RA. The interaction between single mothers' living arrangements and welfare participation. Journal of Policy Analysis and Management 2008;19(1):93–117.

Malpezzi S, Chun GH, Green RK. New place-to-place housing price indexes for U.S. metropolitan areas, and their determinants. Real Estate Economics 1998;26(2):235–274.

McLanahan, S.; Sandefur, G. Growing up with a single parent. Cambridge, MA: Harvard University Press; 1994.

Moffit, RA. The effect of welfare on marriage and fertility: What do we know and what do we need to know?". Moffitt, RA., editor. The effect of welfare on the family and reproductive behavior. National Research Council; 1998.

Moffitt, RA. Department of Economics. Johns Hopkins University; 2005. Welfare benefits database. Available from: http://www.econ.jhu.edu/People/Moffitt/DataSets.html

Nixon L. The effect of child support enforcement on marital dissolution. The Journal of Human Resources 1997;32(1):159–181. doi:10.2307/146244.

O'Flaherty, B. City economics. Cambridge, MA: Harvard University Press; 2005.

Olsen EO. Housing programs for low-income households. 2002 National Bureau of Economic Research. Working paper 8208.

Pong S-L. Sex preference and fertility in peninsular Malaysia. Studies in Family Planning 1994;25(3): 137–148. doi:10.2307/2137940. [PubMed: 7940619]

Powell MA, Parcel TL. Effects of family structure on the earnings attainment process: Differences by gender. Journal of Marriage and the Family 1997;59:419–433. doi:10.2307/353480.

Sigle-Rushton W, McLanahan S. The living arrangements of new unmarried mothers. Demography 2002;39(3):415–433. doi:10.1353/dem.2002.0032. [PubMed: 12205750]

U.S. Department of Housing and Urban Development. A picture of subsidized households—1998. 2005 http://www.huduser.org/datasets/assthsg/statedata98/. Last modified 3/31/05.

U.S. Department of Housing and Urban Development. Low income housing tax credits. 2006 http://www.huduser.org/datasets/lihtc.html. Last modified 10/18/06.

Walker JR. The effect of public policies on recent Swedish fertility behavior. Journal of Population Economics 1995;8:223–251. doi:10.1007/BF00185251.

Willis RJ. A theory of out of wedlock childbearing. The Journal of Political Economy 1999;107(6):533–564. doi:10.1086/250103.

Winkler AE. The impact of housing costs on the living arrangements of single mothers. Journal of Urban Economics 1992;32:388. doi:10.1016/0094–1190(92)90026–H.

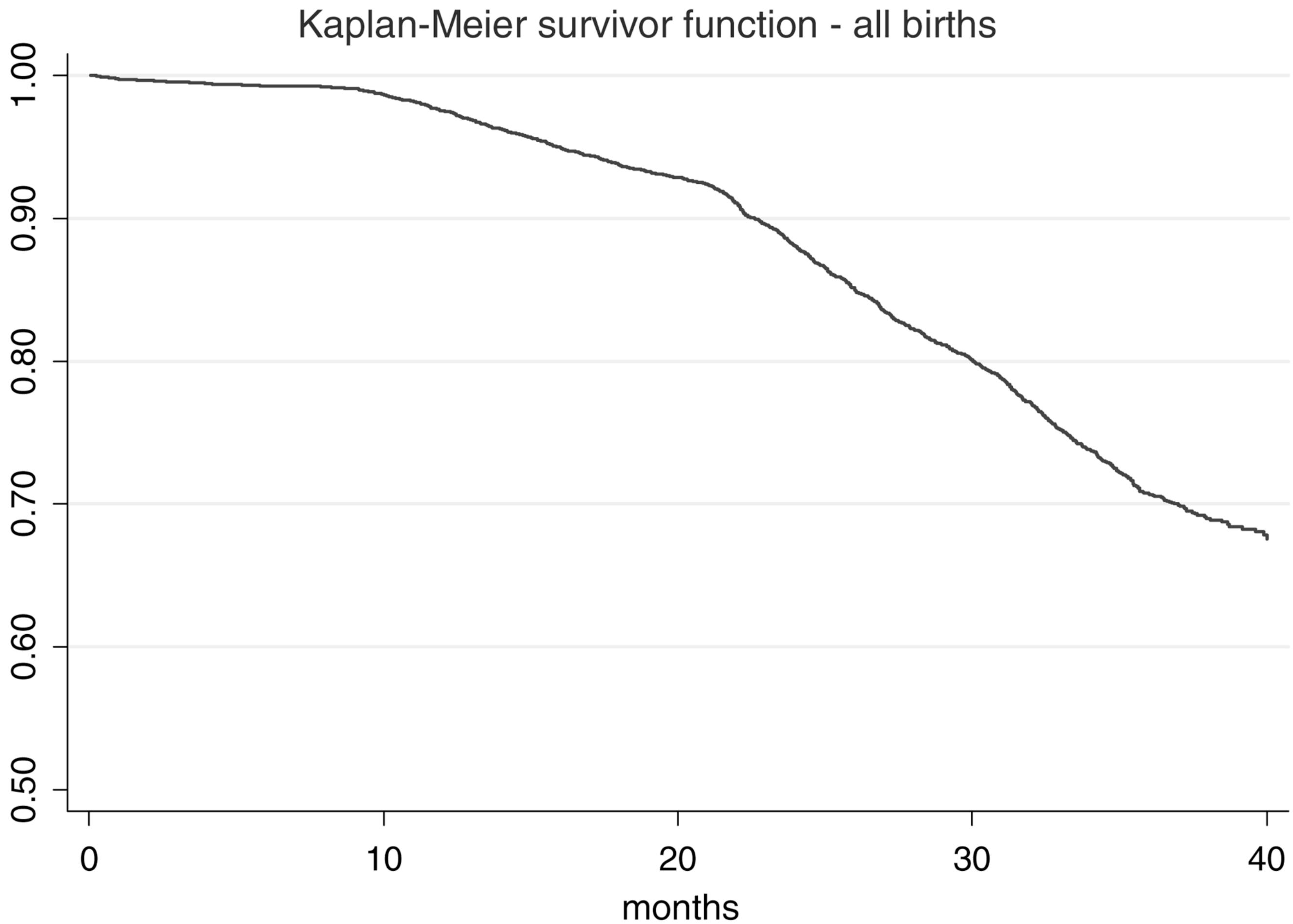

Fig. 1.
Kaplan-Meier survivor function—all births

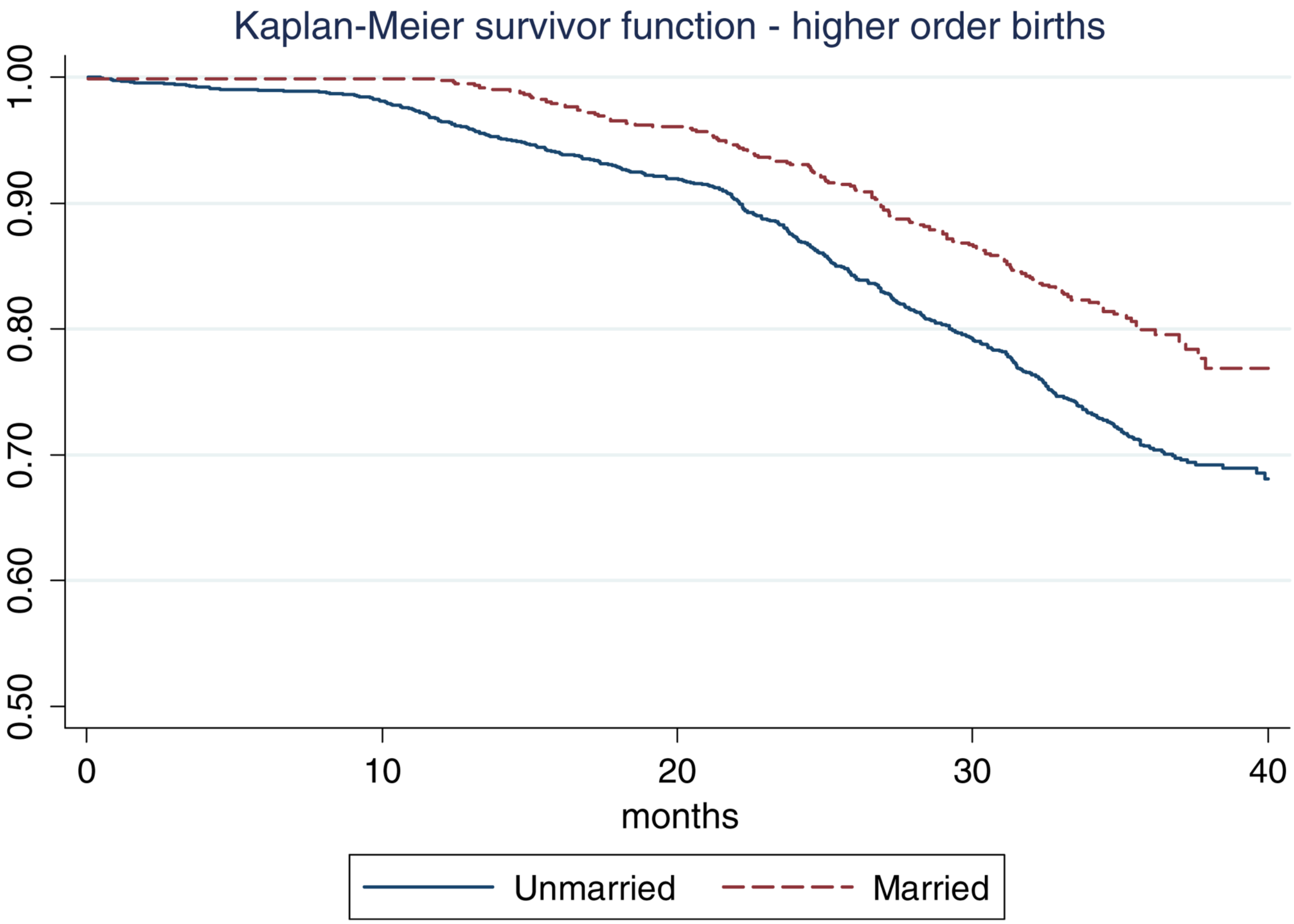

Fig. 2.
Kaplan-Meier survivor function—higher order births—married and unmarried

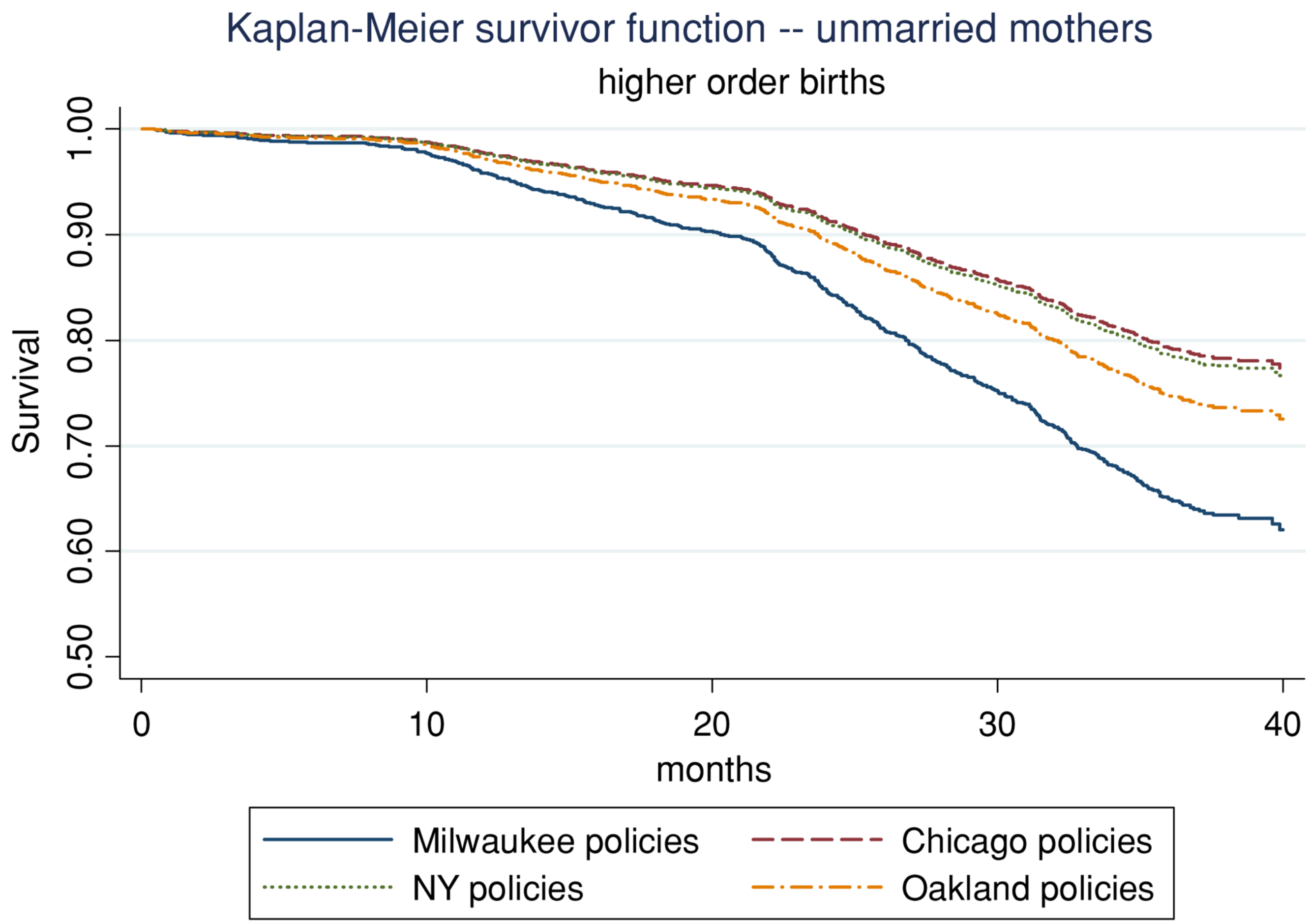

Fig. 3.
Kaplan-Meier survivor function—unmarried mothers: higher order births—city policies

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Table 1

Fertility of fragile families mothers by 3-year follow-up

	% Of sample	% With a subsequent birth	Mean time to birth (in months)
All mothers (N = 4,798)	100	25	29
All married (N = 1,161)	24	25	30
Married, first birth at baseline (N = 412)	9	38	29
Married second or higher birth at baseline (N = 760)	16	18	31
All unmarried (N = 3,637)	76	25	30
Unmarried, first birth at baseline (N = 1,416)	30	25	29
Unmarried, second or higher birth at baseline (N = 2,209)	46	26	29

Table 2

Characteristics of fragile families mothers

	Married no birth N = 867 (%)	**Married birth N = 294 (%)**	**Unmarried no birth N = 2,713 (%)**	**Unmarried birth N = 924 (%)**
Age				
Under 20	2*	7	19*	29
20–24	18	20	40	44
25–29	29	39	22	18
30+	51	35	18	10
Race/ethnicity				
White, non-Hispanic	40	45	14	13
Black, non-Hispanic	25	25	55	54
Hispanic	27	22	28	29
Other	854	7	3	3
Missing information	<1	0	<1	<1
Nativity				
Foreign born	29	25	15*	11
Missing information	<1	<1	<1	<1
Education				
Less than a high school education	17	17	39*	45
High school	21	16	33	33
Some college or more	62	67	27	22
Missing information	0	0	<1	<1
Co-residential status				
Cohabiting	–	–	45	44
Visiting	–	–	45	46
Parents not seeing each other	–	–	9	9
Missing information	–	–	1	1
Number of children at baseline				
Mother has one child	28*	53	38	38
Mother has two children	39	29	31	32
Mother has three or more children	32	18	30	28
Missing information	<1	0	<1	<1
Child characteristics				
Focal child is low birthweight	7	5	13	12
Focal child is a boy	54	55	53	54
Mother worked in the last year	72	78	75	73
Religious observance				
Weekly church attendance	33	36	18	19

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	Married no birth N = 867 (%)	**Married birth N = 294 (%)**	**Unmarried no birth N = 2,713 (%)**	**Unmarried birth N = 924 (%)**
Monthly church attendance	19	16	15	15
Yearly church attendance	23	22	21	19
Hardly/never attend	25	26	46	46
Missing information	<1	<1	<1	<1

Notes: Chi-square tests compare married no birth to married birth (column 2 to column 3); Chi-square tests compare unmarried no birth to unmarried birth (column 4 to column 5); due to rounding all columns do not sum to 100%

$^{*}p<0.05$

Table 3

Birth rates, policy and economic variation by city

	Total N	Number of births	Birth rate	Male unemployment	Log male wages	Subsidized housing	Log house price index	Log TANF + food stamps	Log expenditures child support
Austin, TX	326	109	0.334	0.031	2.143	0.015	11.862	6.457	5.366
Milwaukee, WI	348	109	0.313	0.033	2.276	0.038	11.681	6.817	6.159
Philadelphia, PA	337	105	0.312	0.055	2.246	0.039	11.803	6.596	5.974
Indianapolis, IN	325	96	0.295	0.019	2.177	0.027	11.402	6.435	5.182
Baltimore, MD	338	93	0.275	0.033	2.297	0.080	11.777	6.531	5.999
Corpus Christi, TX	331	86	0.260	0.037	2.226	0.022	11.201	6.277	5.407
Toledo, OH	101	26	0.257	0.044	2.264	0.036	11.463	6.557	6.349
Norfolk, VA	99	25	0.253	0.027	2.312	0.038	11.529	6.439	5.602
Pittsburgh, PA	100	25	0.250	0.036	2.252	0.151	11.369	6.604	6.040
Richmond, VA	327	81	0.248	0.009	2.386	0.074	11.478	6.430	5.811
Oakland, CA	330	80	0.242	0.054	2.253	0.034	12.509	6.988	5.966
San Antonio, TX	100	24	0.240	0.037	2.226	0.022	11.434	6.277	5.407
Detroit, MI	327	78	0.239	0.042	2.325	0.070	11.778	6.650	6.061
Newark, NJ	342	78	0.228	0.050	2.478	0.142	12.293	6.609	6.219
Nashville, TN	102	23	0.225	0.032	2.223	0.044	11.564	6.254	5.352
Boston, MA	99	22	0.222	0.021	2.432	0.142	12.469	6.749	5.697
Chicago, IL	155	29	0.187	0.042	2.356	0.046	11.972	6.532	5.635
San Jose, CA	327	60	0.183	0.043	2.315	0.021	12.955	6.773	6.151
New York, NY	384	69	0.180	0.038	2.342	0.041	12.599	6.758	5.680
Observations	4,798	1,218							

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Table 4

Risk of subsequent fertility among families with at least two children at baseline—city fixed effects

	Higher birth—married		Chi-square (Prob > chi-sq)	Higher birth—unmarried	
	Hazard ratio	Coeff. (SE)		Hazard ratio	Coeff. (SE)
Age					
Mother is 20–24 (ref)					
Mother is under 20	2.269**	0.819**(2.310)	0.92 (0.34)	1.567***	0.449***(3.680)
25–29	1.107	0.101 (0.370)	1.51 (0.22)	0.712***	−0.340*** (2.920)
30 or older	0.497**	−0.699* (2.510)	0.01 (0.92)	0.515***	−0.664*** (4.670)
Race/ethnicity					
Black, non-Hispanic (ref)					
White, non-Hispanic	0.752	−0.285 (1.090)	1.48 (0.22)	1.134	0.126 (0.760)
Hispanic	0.985	−0.015 (0.050)	0.59 (0.44)	1.252*	0.225* (1.690)
Other	0.426	−0.854 (1.460)	0.07 (0.79)	1.813**	0.595** (2.340)
Foreign born	0.927	−0.076 (0.260)	0.00 (0.96)	0.911	−0.093 (0.540)
Mother's education					
College education (ref)					
Less than high school education	0.901	−0.105 (0.360)	0.08 (0.78)	0.979	−0.021 (0.160)
High school	0.668	−0.403 (1.630)	2.19 (0.139)	1.002	0.002 (0.020)
Co-residential status					
Parents not seeing each other (ref)					
Cohabiting	–	–	–	0.846	−0.168 (1.130)
Visiting	–	–	–	0.899	−0.107 (0.720)
Mother has three or more children (ref)					
Mother has two children	1.313	0.272 (1.360)	3.07 (0.08)	0.846*	−0.167* (1.780)
Focal child low birthweight	0.995	−0.005 (0.010)	0.00 (1.00)	1.000	0.000 (0.000)
Focal child is a boy	1.099	0.094 (0.530)	0.12 (0.73)	1.017	0.017 (0.200)
Mother worked in the past year	1.107	0.102(0.470)	0.03 (0.87)	1.155	0.144 (1.430)

	Higher birth—married		Chi-square (Prob > chi-sq)	Higher birth—unmarried	
	Hazard ratio	Coeff. (SE)		Hazard ratio	Coeff. (SE)
Religious observance					
Hardly/never attend (ref)					
Weekly church attendance	0.99	−0.010 (0.040)	0.54 (0.46)	1.172	0.159 (1.360)
Monthly church attendance	0.728	−0.317 (1.100)	0.47 (0.49)	0.887	−0.120 (0.930)
Yearly church attendance	0.936	−0.066 (0.250)	0.00 (0.95)	0.918	−0.086 (0.720)
City fixed effects					
Milwaukee, WI (ref)					
Oakland, CA	0.987	−0.013 (0.030)	1.15 (0.28)	0.566***	−0.568*** (2.600)
Austin, TX	0.874	−0.135 (0.280)	0.02 (0.88)	0.802	−0.221 (1.080)
Baltimore, MD	1.16	0.148 (0.330)	0.32 (0.57)	0.857	−0.155 (0.770)
Detroit, MI	0.626	−0.469 (0.950)	0.10 (0.75)	0.752	−0.285 (1.370)
Newark, NJ	0.627	−0.467 (0.920)	0.01 (0.94)	0.657*	−0.421* (1.870)
Philadelphia, PA	1.181	0.166(0.360)	0.32 (0.57)	0.926	−0.076 (0.390)
Richmond, VA	0.967	−0.033 (0.070)	0.07 (0.79)	0.838	−0.177 (0.840)
Corpus Christi, TX	0.323**	−1.131** (1.980)	1.68 (0.19)	0.684*	−0.380* (1.720)
Indianapolis, IN	0.580	−0.544 (1.030)	0.41 (0.52)	0.888	−0.119 (0.580)
New York, NY	0.807	−0.214 (0.450)	0.78 (0.38)	0.450***	−0.799*** (2.900)
San Jose, CA	0.666	−0.406 (0.730)	0.00 (0.99)	0.630*	−0.462* (1.710)
Boston, MA	0.258	−1.356(1.260)	0.01 (0.93)	1.152	0.142 (0.440)
Nashville, TN	0.419	−0.87 (1.100)	0.00 (0.99)	0.516*	-0.661* (1.660)
Chicago, IL	0.732	−0.311 (0.530)	0.37 (0.54)	0.495**	−0.703** (2.200)
Toledo, OH	0.596	−0.518 (0.650)	0.00 (0.98)	0.803	−0.220 (0.690)
San Antonio, TX	0.597	−0.516 (0.620)	0.00 (0.99)	0.531*	−0.633* (1.850)
Pittsburgh, PA	1.339	0.292 (0.530)	0.77 (0.38)	0.719	−0.330 (0.850)
Norfolk, VA	1.022	0.021 (0.040)	0.37 (0.55)	0.689	−0.373 (1.030)
Observations	760			2,209	
Log Likelihood	−	847.730			−4052.488

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	Higher birth—married		Chi-square (Prob > chi-sq)	Higher birth—unmarried	
	Hazard ratio	Coeff. (SE)		Hazard ratio	Coeff. (SE)
Chi-square (df)	59.76 (34)			98.93 (36)	
Chi-square (df), joint sig. city fx	14.25 (17)			18.49 (17)	
Prob > chi-square	0.65			0.36	

Note: Wald Test for difference in coefficients between models

*$p < 0.10$,

**$p < 0.05$,

***$p < 0.01$

Table 5

Unmarried mothers' risk of higher order birth—policy effects

	Hazard ratio	Coeff. (SE)
Mean male unemployment rate	0.242**	−1.419** (2.64)
Mean male log wages	0.060*	−2.819* (2.48)
Availability of subsidized housing	1.177	0.163 (0.81)
Housing costs	1.009	0.009 (0.05)
Housing costs*availability of subsidized housing	2.018*	0.702* (2.13)
TANF/FS grant amount	0.987	0.013 (0.03)
Total expenditures on child support enforcement	2.026**	0.706** (2.58)
Observations	2209	
Log likelihood	−4056.060	
Chi-square (df)	94.75 (26)	
Chi-square (df), joint sig. policy variables	14.73 (7)	
Prob > chi-square	0.04	

Note: Model includes the following covariates: mother is under 20, 25–29, 30 or older, White, non-Hispanic, Hispanic, other race, foreign born, less than high school, high school, cohabiting, visiting, two children, low birthweight, focal child is a boy, worked in past year, attends church weekly, attends church monthly and attends church yearly

* $p<0.05$,

** $p<0.01$

Table 6

Unmarried mothers' risk of second birth—policy effects

	Hazard ratio	Coeff. (SE)
Mean male unemployment rate	0.778	−0.252 (0.380)
Mean male log wages	0.350	−1.051 (0.740)
Availability of subsidized housing	0.926	−0.077 (0.310)
Housing costs	0.690	−0.371 (1.620)
Housing costs*availability of subsidized housing	2.434*	0.889* (2.190)
TANF/FS grant amount	1.695	0.528 (0.970)
Total expenditures on child support enforcement	1.047	0.046 (0.130)
Observations	1,416	
Log likelihood	−2423.888	
Chi-square (df)	55.51 (25)	
Chi-square (df), joint sig. policy variables	20.64 (7)	
Prob > chi-square	0.00	

* $p< 0.05$

Note: Model includes the following variables: mother is under 20, 25–29, 30 or older, White, non-Hispanic, Hispanic, other race, foreign born, less than high school, high school, cohabiting, visiting, low birthweight, focal child is a boy, worked in past year, attends church weekly, attends church monthly and attends church yearly

Table 7

Simulation results—policies and changes in unmarried mothers higher order fertility

	New York % without a birth)	Milwaukee (% without a birth)	Difference
Housing costs	81	75	+6
Total expenditures on child support enforcement	81	75	+6
Mean male log wages	79	75	+4
Availability of subsidized housing	78	75	+3
Housing costs* availability of subsidized housing	69	75	−6
Mean male unemployment rate	77	75	+2
TANF/FS grant amount	77	75	+2