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Race/Ethnic Differences in Fertility Behaviors: Early Childbearing and Number of Children

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Using the GSS 1994-2008, I analyze racial/ethnic differences in fertility behaviors among Whites, African Americans, Mexicans, other Hispanics, and Asians. I focus on two key elements of family formation, age of initiation of childbearing, and age-specific cumulative fertility, estimated by the number of children ever born. I use survival analysis and discrete-time Cox regression to analyze the age at time of first birth. A Poisson regression is used to compare the number of children ever born. The analysis reveals that racial differences in fertility behaviors are greatly explained by the timing of motherhood initiation. African Americans and Hispanics have significantly greater risks of having a first child at younger ages compared to Whites. However, the racial/ethnic differences in the total number of children are small after we control for women's age at time of first birth. Structural opportunities seem to pattern childbearing behaviors. This analysis supports the racial stratification perspective.

Patterns of family formation in the U.S. have been changing in the last few decades at the same time that recent waves of immigration have altered the ethnic composition of the U.S. population (Casper & Bianchi, 2002). Recently-arrived Hispanic and Asian groups are growing at faster rates than previously established ethnic groups, increasing their share of the total population. The two main sources of this growth are the immigrant influx, and natural growth or reproduction by birth. As previous studies have shown, family behaviors vary greatly by racial/ethnic groups (Frank & Heuveline, 2005; Landale, Schoen & Daniels, 2010).

Increasingly, Americans are deciding to retreat from marriage and remain childless. The current trends include a decline in marriage rates, a decrease in marital fertility, increases in cohabitation, and an increase in non-marital childbearing (Casper & Bianchi, 2002; Landale, Oropesa, & Bradatan, 2006). However, not all racial minority groups seem to follow the same patterns of family formation found among the White middle-class. African Americans and Hispanics exhibit higher rates of early childbearing and fertility than those found among White females. Some scholars emphasize that family structure and family formation have a major effect on the socioeconomic outcomes and the welfare of successive generations (Biblarz & Raftery, 1999; Mare & Maralani, 2006; Musick & Mare, 2004; Rumbaut, 2005).

However, most of the previous research has relied on data collected prior to 1995 (Oropesa & Landale, 2004; Wildsmith & Raley, 2006), providing little insight into today's new ethnic groups and the extent to which their childbearing behaviors are comparable to the patterns found in the White majority U.S. population. Moreover, because of their relatively recent arrival, Asian groups have not been included in most studies about fertility. The purpose of this study is to evaluate the effect of race/ethnicity on fertility behaviors. I focus on two key elements of family formation, the age of initiation of childbearing, and age-specific cumulative fertility, estimated by the number of children ever born.

The literature in fertility has neglected the important role of childbearing initiation to explain fertility differentials. Most studies on fertility either focus on age at time of first birth or on total number of children. In this paper, I use both variables to compare the differences in the risk of having a first child and age-specific fertility by racial/ethnic group. I find that the difference between Whites and most of the other minority groups in the age at first birth are large and significant, whereas the differences in number of children beyond the first one are smaller and, in some cases, not statistically significant.

Most of the previous research on fertility has relied on OLS regression. OLS regression methods pose several limitations to fertility analysis, such as the inability to account for censored data or time-varying covariates (Allison, 2010; Raftery, Aghajanian, & Kahn, 1996). Survival analysis represents a more suitable approach to analyze fertility, as it allows for the inclusion of censored cases, such as childless women, who are also at risk of having children. In this study I use recent data to evaluate the effect of race/ethnicity on fertility behaviors among different ethnic groups taking into account childless women (censored cases) and the age of the mother at first birth, controlling for education and other factors. I analyze two outcome variables of family formation, the age of initiation of childbearing and age-specific cumulative fertility, estimated by the number of children ever born. In the first part, I use survival analysis to compare the age-specific risk of having a first child by race/ethnicity. Then, using a Poisson regression, I show that some of the fertility differential across racial/ethnic groups is greatly explained by the differences in mother's age at first birth.

This study uses the General Social Survey (GSS) 1994-2008, a representative national sample collected via interviews administered by the NORC using a standard questionnaire. I restrict my analysis to women ages 20 to 49. The dataset includes a variable for the age when mothers had their first child, and a variable for the total number of children ever born for each respondent, and it allows for the identification of different ethnic groups including White and African American females, as well as Hispanic and Asian origin populations.

Background

Post 1965 immigrants are changing the demographic composition of the U.S. population as recently-arrived Hispanic and Asian groups are growing at faster rates than previously established ethnic groups, increasing their share of the total population. The new immigrant groups present different cultural traits and demographic profiles than the ones exhibited by earlier European immigrants. Moreover, the economic and social context of reception for immigrants have changed raising concerns about whether these new ethnic groups and their children will be able to assimilate into the U.S. mainstream and adopt the norms and behaviors found in this society.

Explaining Racial/Ethnic Differentials: Cultural vs. Structural Explanations

Most studies about racial/ethnic differentials in fertility behaviors resort to two types of explanations to account for these differences. One of the explanations attributes fertility variation to differences in attitudes and cultural values. Others attribute differences to structural factors such as socioeconomic background and societal context. More recent approaches try to integrate both accounts.

According to the cultural explanation, Hispanics hold a set of cultural values that emphasizes the importance of the family, and that encourages family formation and childbearing. This set of cultural values is called "familism." Familism and religious beliefs have been used to explain patterns of family formation, such as high marriage rates and early fertility among Mexican and other Hispanic origin groups (Landale, Schoen & Daniels, 2010; Oropesa & Landale, 2004). However, while this explanation could account for the higher marriage rates found among Hispanics compared to African Americans, it does not explain why these two groups exhibit similar patterns of early childbearing initiation. Moreover, other scholars also note that family formation patterns and cultural values have also been changing in immigrants' countries of origin, such as the recent trends of declining fertility in Latin American countries (Landale & Oropesa, 2007).

The classical assimilation theory proposed by Gordon (1964) considers assimilation as a gradual process across generations by which new ethnic groups shed their cultural traits and acquire behavioral patterns that resemble those found in the host society. According to this vision the original cultural values and norms are replaced by the ones present in the dominant group, embodied by the white middle-class. Acculturation was considered the first step in the process, and a requirement for socioeconomic mobility (Alba & Nee, 2005; Gordon, 1964). This theory has been criticized for assuming that assimilation into the white middle-class is the only possible outcome of successful assimilation and for ignoring the contextual factors that affect the incorporation process. The theory of segmented assimilation proposed by Alejandro Portes and Min Zhou also accounts for some of the limitations of classical assimilation theory and integrates structural explanations. Portes and Zhou (1993) consider alternative patterns of assimilation depending on immigrants' human and social capital, context of reception, and place of settlement upon arrival (Portes & Zhou, 1993). According to this theory, disadvantaged immigrants, such as some Hispanic groups who settle in poor neighborhoods, may adopt the habits of their disadvantaged neighbors and incorporate instead into a lower class.

Following the tenets of assimilation theory, recent studies have noticed that familism values decrease with time in the U.S. as new cohorts adopt the individualistic ideals of Western societies (Lloyd, 2006; Telles & Ortiz, 2008). Thus, although cultural values can be used to explain differentials in family formation patterns, they should not be considered to be fixed, but constantly changing.

An alternative explanation, however, considers that structural factors such as socioeconomic status, including education, availability of economic resources, and the characteristics of the family of origin, exert a great influence on subsequent family formation behaviors. Using in-depth interviews some scholars have found that minority groups who are skeptical about the value of education may pressure youngsters to start families at an early age (Kasinitz, Mollenkopf, Waters & Holdaway, 2008). Edin and Kefalas (2005) affirmed that women who live in low-income neighborhoods start bearing children at an early age and exhibit high fertility rates due to a series of factors associated with the limited educational and economic opportunities available to themselves and their partners, regardless of their ethnic background.

In their study about fertility differentials, Frank and Heuveline (2005) also emphasized the importance of the U.S. social context in shaping fertility outcomes. Their study supports the racial stratification perspective that poses that American society, which is racially stratified, provides different opportunity costs to some minority groups in educational opportunities, career expectations, and other options for socioeconomic mobility that influence fertility behaviors (See also Telles & Ortiz, 2008).

However, most of the literature on stratification and immigrant assimilation has focused on socioeconomic outcomes such as educational attainment, income, homeownership, and

English proficiency (Myers, 2007), but less attention has been paid to patterns of family formation. Other studies have mainly treated family formation as an explanatory factor of SES. In this paper I explore family formation patterns as an outcome variable. In recent times, family demographers and other scholars have become more interested in predicting whether the new ethnic groups will maintain the childbearing behaviors prevalent in their country of origin, and how these behaviors may affect and be affected by the patterns already in place in the U.S.

During the last decades, mainstream patterns of family formation in the U.S. have drastically changed. The current trends include a decline in marriage rates, a decrease in marital fertility, increases in cohabitation, and an increase in non-marital childbearing (Casper & Bianchi, 2002; Landale, Oropesa, & Bradatan, 2006). Nonetheless, childbearing behaviors among some Hispanic groups do not seem to follow the patterns found among the white majority. In contrast, Hispanics exhibit high rates of early childbearing that more closely resemble those of the African American population (Frank & Heuveline, 2005; Telles & Ortiz, 2008; Wildsmith & Raley, 2006). Based on these theories, the present study compares White, African American, Hispanic, and Asian women, and evaluates how family formation behaviors vary across these groups.

Initiation of Motherhood and Early Childbearing

The recent trends in the U.S. indicate that more middle-class women are delaying childbearing and many are remaining childless (Casper & Bianchi, 2002). The age at which a woman starts bearing children is an important factor in explaining socioeconomic mobility because it can alter the life course and influence the life chances of women. In a recent study, Rumbaut concluded that early childbearing among Hispanic women significantly thwarts their socioeconomic mobility as it becomes an obstacle for educational and occupational mobility, perpetuating the cycle of economic disadvantage (Landale et al., 2009; Rumbaut, 2005).

Early childbearing has proved to be highly correlated to education, but the directionality of causation is not always clear. A possible explanation is that women intentionally delay marriage and childbearing until they achieve the level of education they expect, in which case higher education would be causing delays in childbearing (Mare & Maralani, 2006). Nonetheless, some scholars note that when the prospect of achieving higher education are dim, education does not constitute a strong incentive to delay childbearing, as it is in the case of women in low-income neighborhoods (Edin & Kefalas, 2005). In other words, early childbearing carries high opportunity costs, or so it is perceived, for middle-class and upper-class females, but not for poor women. An alternative possibility is that some women are prevented from completing their education when they become pregnant, causing them to dropout of high school or college, or to forgo higher education altogether. In this case, early pregnancy would be preventing higher educational attainments (Rumbaut, 2005). A third possibility is that both education and childbearing behaviors depend on common exogenous causes, such as future expectations, cultural values, and/or opportunities. (Edin & Kefalas, 2005; Kasinitz et al., 2008; Landale et al., 2009).

In this analysis I find a significant association between women's education and fertility behaviors. However, because the GSS does not provide data on educational attainment prior to the onset of pregnancy, there is a potential problem of endogeneity to evaluate causality or effect directionality. Without this information the results could be misleading since some females continue their education after they have a first child. However, it is possible to control for education when analyzing fertility behaviors, and this is what I do in the analysis in this study. The evaluation of the third alternative requires additional analysis that is not conducted in the present paper because of high levels of missing data on the exogenous variables.

Fertility Trends and Racial/Ethnic Differentials

Demographers have long documented the current trend of declining fertility rates prevalent in most countries (Bongaarts & Potter, 1983). Declining fertility has raised concerns about the potential contraction of the labor force in developed countries (Myers, 2007). In addition, fertility rates also have direct implications for the future demographic composition of the population. As many studies reveal, Hispanic groups exhibit the fastest growth rates of all ethnic groups in the U.S. in great part due to their relatively high fertility rates, and to a lower extent, to the continuous influx of Hispanic immigrants (Frank & Heuveline, 2005; Telles & Ortiz, 2008). Asian groups in the U.S. have also been increasing their share of the total population; however, despite the diversity of this demographic group, Asians tend to delay marriage and childbearing, and they are also characterized by relatively low fertility rates (Hwang & Saenz, 1997).

Class also matters as it conditions educational outcomes and other structural opportunities. Poor married and unmarried women exhibit relatively high fertility, regardless of their ethnic background (Edin & Kefalas, 2005; Musick & Mare, 2004). Kasinitz et al. (2008) state that "In general, children of more educated, wealthier families tend to postpone marriage and parenting, whereas single parenthood and teenage pregnancy are most common among women from poor families of all racial and ethnic backgrounds" (p. 207). Although there is debate on what is the best way to measure socioeconomic status (SES), education and income are usually among the preferred proxies. Thus in addition to education, I control for family income in this study.

I analyze the racial/ethnic differentials on the motherhood initiation, measured by the women's age at first birth, and age-specific cumulative fertility, measured by number of children ever born. Following the literature about ethnic differences in fertility behaviors, I expect to find a significant racial effect on both outcome variables. As this analysis will show, I found evidence to support this hypothesis; however, I also found that a significant part of the racial effect on fertility can be explained by the age when women begin bearing children.

Data and Methods

To conduct this analysis I use the General Social Survey (GSS) 1994-2008. Before 1994 the GSS did not include a variable for mother's age at first birth, thus I only use waves from 1994 to 2008 in which the variables of interest are available. The GSS includes individuals ages 18 and over. However, I restrict my analysis to female respondents ages 20 to 49 in order to generate evenly spaced age-group cohorts, following the standard 5-year age groups for fertility analysis: 20-24, 25-29, 30-34, 35-39, 40-44, and 45-49 (Preston, Heuveline, & Guillot, 2001). The total sample size was 7,763; however, 531 cases were eliminated due to missing values on some of the main variables[1], leaving a total sample size of 7,232.

As a proxy to measure fertility, I use parity or age-specific cumulative fertility measured by the total number of biological children ever born alive to female respondents. This variable has been used in previous research as it constitutes a very good proxy to study fertility (Frank & Heuveline, 2005; Preston, et al., 2001; Raftery, et al., 1996). Because *children ever born* is not a normally distributed variable, I use a Poisson regression instead of a OLS regression for this part of the analysis. Since fertility is highly dependent on age, all the models control for time exposure to fertility measured in years of age. In addition, I include 5-age categories to account for age-cohort differences; the 20 to 24 age category is used as the referent group.

The variables *race* and *ethnic* in the GSS were used to identify racial/ethnic groups. Race/ethnicity is included as a set of mutually exclusive dummy variables for African Americans, Mexicans, other Hispanics, Asians, and other races. Whites are the referent group in all the models. Education is used as a control variable and is introduced as dummy variables

[1] Individuals with missing values on age, education, number of children, marital and employment status, number of siblings, and nuclear family upbringing were dropped from the sample.

for those with less than high school, high school (referent), some college, and college education. Based on the literature about fertility, I control for original family structure by including a dummy variable for those who lived with both parents at age 16 (=1, 0 otherwise). I also control for number of siblings, which is introduced as a continuous variable. The variable *born* (was the respondent born in the U.S.?) is used to identify native-born and immigrant women.

To control for income I used the variable *realinc* which reports family income per year adjusted by inflation to avoid problems of noncomparability due to changes in real value over time. Income is reported in constant dollars using 1986 as the year base. For the last survey year, 2008, in which the variable *realinc* is missing, I used the variable *income06* in which income is measured by intervals. I calculated the midpoints of each interval and adjusted the midpoint income for inflation based on the consumer price index. The midpoint of the last open-ended interval was calculated by adding 25% to the lower limit in that interval, following Hout's calculations (Hout, 2004). For the regression analyses, real income is divided by 10,000 to obtain the effect per every $10,000 increase in yearly family income. Cases with missing values in income were imputed using the mean values specific to each race group (a total of 656 cases, representing 9.07% of the sample, were imputed). A dummy variable for imputed income values was tested in the models, and it was not significant.

To control for religious affiliation I conducted various preliminary analysis based on the literature. Unfortunately, I could not separate Evangelicals or born-again Christians from Protestants and Catholics using the GSS. I used the variable *relig16* which reports the religion in which the respondents were raised. When *relig16* was missing I used the variable *relig*, (respondent's religious affiliation at the time of the survey) to impute values on *relig16* (only 31 cases were imputed). Following previous studies I introduce a dummy variable for Catholics (Oropesa, 1996).

The GSS does not ask individuals for their age at first marriage anymore, so I could not control for age at marriage in this study. Age at marriage used to be an important determinant of fertility (Bongaarts & Potter, 1983). However, increases in cohabitation and nonmarital childbearing, have limited the strength of age at marriage to predict fertility. Nonetheless, in the second part of the analysis, I introduce a dummy-variable to control for marital status (1=married, 0 otherwise). Unfortunately, the GSS does not include data on cohabiting status, so I could not control for cohabitation. I also control for employment status (employed =1, 0 otherwise) in the fertility analysis. Year of survey was introduced in all the regression models as a continuous variable to control for period effects.

In the first part, I conduct an exploratory survival analysis on the age at time of first birth by ethnic groups. Survival analysis is the most suitable method to analyze age at first birth since this method allows for the inclusion of censored cases (childless women) in the analysis. Censored cases are those in which the observation period is terminated before the event (having a first child) occurs, if it ever occurs. All women are considered at risk of having a birth until they have a first child, or until the time when the survey was conducted if they did not have a child by that time. Then I use Cox regression to analyze the age-specific risk (or hazard rate) of having a first child while controlling for other factors. The results show hazard ratios of the age-specific probability of childbearing initiation for selected ethnic groups.

Table 1 provides descriptive statistics for all the variables by race-ethnicity. The total sample size is 7,232 of which 66% is white, and the rest are minority groups. For descriptive purposes only, number of children ever born is displayed in four categories (0, 1, 2, 3 or more children). It is worth noting that Asians constitute the group with the largest percentage of childless women (36%), followed by Whites (32%), for the rest of the groups the percentage of childless women vary from 23% to 18%. We also observe great variability on the age at first birth, with Asians and Whites reporting higher means for age at first birth, 26.3 and 23.7 respectively, for the other minority groups, the mean age at first birth is around 21 years of age.

Table 1. Percent Distribution of all Variables by Race/Ethnicity for Women ages 20 to 49. GSS, 1994-2008 (N = 7,232).

		White	Black	Mexican	Other Hispanic	Asian	Other	Total
	N	4,784	1,190	375	232	223	428	7,232
	%	66.2	16.5	5.2	3.2	3.1	5.9	100.0
Mean Age		35.7	34.7	32.8	34.0	34.2	34.4	35.2
Foreign Born	%	4.2	6.6	38.1	53.9	70.4	8.2	10.2
Number of children	%							
0		32.4	19.3	23.5	18.1	36.3	19.4	28.7
1		20.1	24.0	16.8	19.0	19.7	17.3	20.4
2		27.8	25.6	22.9	27.6	27.4	33.4	27.5
3 or more		19.6	31.2	36.8	35.3	16.6	29.9	23.4
Mean age at first birth		23.7	20.6	21.3	21.2	26.3	21.2	22.8
Education								
Less than High School		6.5	16.3	27.2	23.7	6.7	15.4	10.3
High School		52.5	59.0	54.1	53.5	35.4	62.2	53.7
Some College		10.1	9.2	7.7	7.3	9.4	9.8	9.7
College+		31.0	15.5	10.9	15.5	48.4	12.6	26.4
Mean No. of Siblings		3.2	5.0	5.1	4.5	3.5	4.1	3.7
Nuclear Fam. 16		69.4	43.3	65.1	53.0	79.4	53.0	63.7
Employed		72.9	69.8	62.7	65.9	71.3	63.1	71.0
Median Family Income[1]		30,289	16,363	18,947	18,947	37,895	20,065	24,782
Married		54.3	25.1	51.5	34.1	60.1	49.5	48.6
Religious Upbringing								
Protestant		54.2	80.7	13.6	12.1	19.3	64.0	54.6
Catholic		32.2	8.7	82.9	79.7	34.1	19.2	31.8
None		8.3	7.3	2.1	6.5	10.3	7.9	7.8
Other		5.3	3.3	1.3	1.7	36.3	8.9	5.8

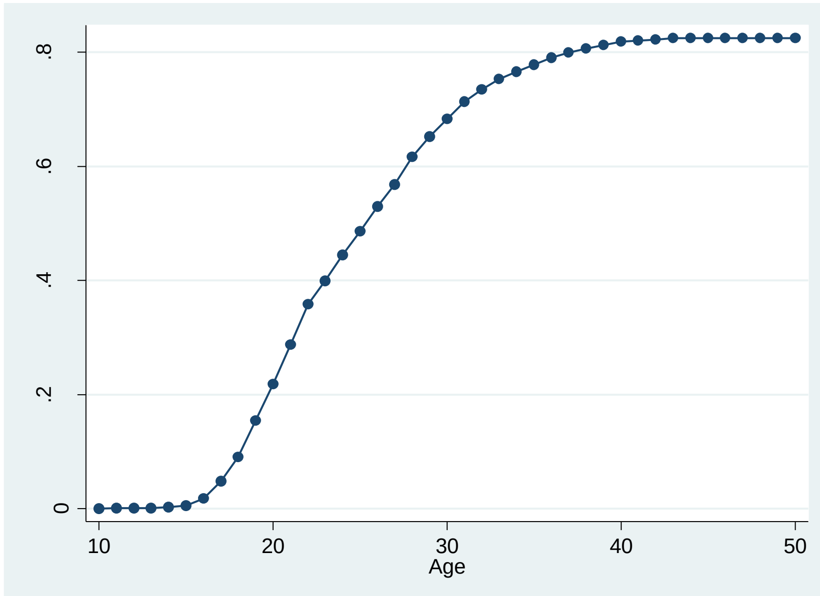
[1]Real income adjusted for inflation. In constant dollars, base = 1986.

It is also worth noting that educational attainment also varies greatly by racial/ethnic groups. Asians report the highest education with more than 48% with college degrees and less than 7% with less than high school education. Whites constitute the next most educated group with 31% with college degrees and also less than 7% with less than high school education. The rest of minority groups follow after a gap, with Mexicans reporting the lowest educational attainment. It is also important to note that African Americans report the lowest percentage of nuclear family upbringing, with only 43% of them having lived with both parents at age 16. They also report the lowest median family income and the lowest percentage of married women with only 25% of them married at the time of the survey. Finally, 81% of African Americans reported being raised as Protestants, however based on the literature, a large number of them may be Evangelical Protestant, a group that is not distinguished among the variables reporting religious affiliation in the GSS.

Preliminary Results

I use discrete-time survival analysis to explore patterns of age at time of first birth. The hazard (or failure) function estimates the age-specific probability of having a first child (the failure event) conditional upon the subject not having a failure event before that age. I begin comparing the failure functions for age at first birth. Figure 2 shows the cumulative failure function for the entire sample without any control variables. The shape of the curve reveals that the proportion of women having a first child rapidly increases from the late teens until the late 20's, then the cumulative proportion of mothers increases at a decreasing rate, and remains relatively constant after age 40. In the sample provided by the GSS, 71.4% of women have had at least one child. The mean age at first birth for those who had children is 22.8 years, and the median age is 22 years, meaning that half of the women in the sample who had children experienced their first birth by age 22.

Figure 2. Cumulative Failure Function: Age at Birth of First Child



Childless women are included in the estimation of the failure function. They are considered to be at risk of having a child until the time of the survey, thus childless women are part of the denominator in the calculations of the probability of having a first birth until the age they had at the time of the survey. At that age they are considered censored cases and are dropped from the population at risk. Similarly, once a woman have a child, she is not longer at risk of having a first birth again, thus she is also removed from the population at risk after she reports having had a child.

The hazard rate measures the risk or intensity at which an event occurs at a particular time, it aims "to quantify the instantaneous risk that an event will occur at time *t*." (Allison, 2010, p.16). A hazard rate of .1 at age 20 means that, were that rate to continue for that entire year, we would expect to see .1 first births per women at age 20. Hazard rates are associated with probabilities, but similar to odd ratios, hazards are reported as ratios and, thus, they can be higher than 1. Figure 3 shows the smoothed hazard rate estimates for age at first birth by race/ethnicity without any control variables. The lower line in the middle represents the hazard function for Whites. Compared to Whites, all other minority groups, except for Asians, exhibit a higher risk of having a first child that peaks during the early 20's. By contrast the risk of having a first child for Asians peaks in the early 30's. The Log-rank and Wilcoxon tests reveal that the differences in hazard rates by ethnic groups are highly statistically significant (p< .001).

Figure 3. Smoothed Hazard Rates by Racial/Ethnic Group: Age at Time of First Birth

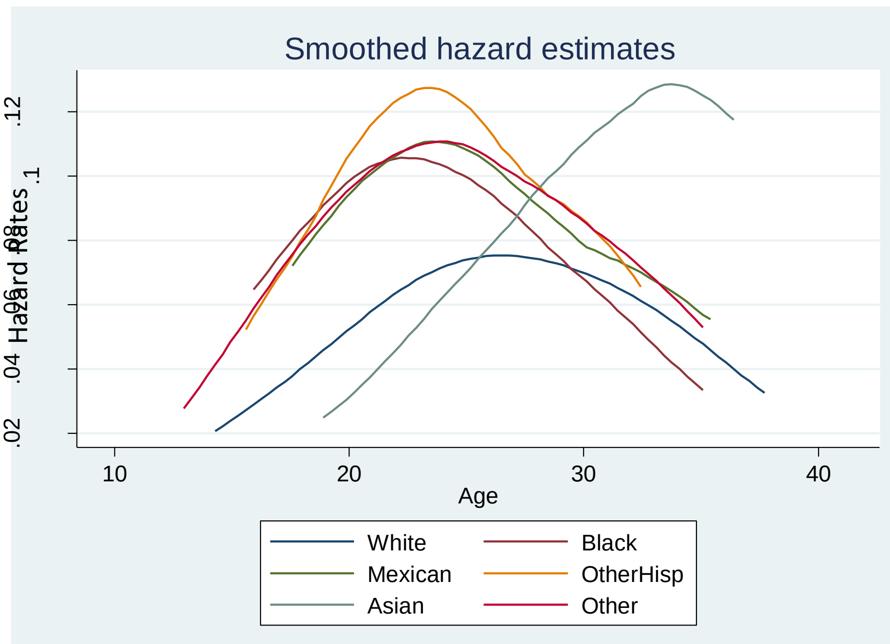
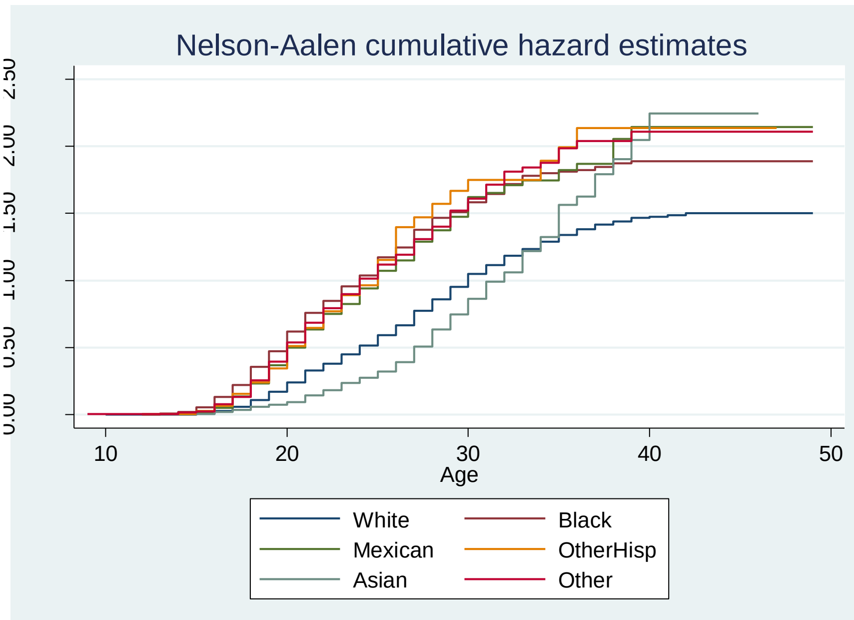


Figure 4 shows the cumulative hazard by racial group. It becomes clear in this graph that Asians and Whites, represented by the two lines at the bottom, exhibit significantly lower hazard rates of having a first birth compared to all other minority groups. Whereas the cumulative hazard functions for all other minority groups are higher and very similar to each other. It is important to remind the reader to be careful when interpreting the hazard estimates at the upper tail. Since females with higher fecundability[2] have children at younger ages, they are dropped from the population at risk at earlier times, leaving the ones with lower

[2] Fecundability refers to the biological component of fertility and it is the potential reproductive capacity of an individual (Preston, et al., 2001).

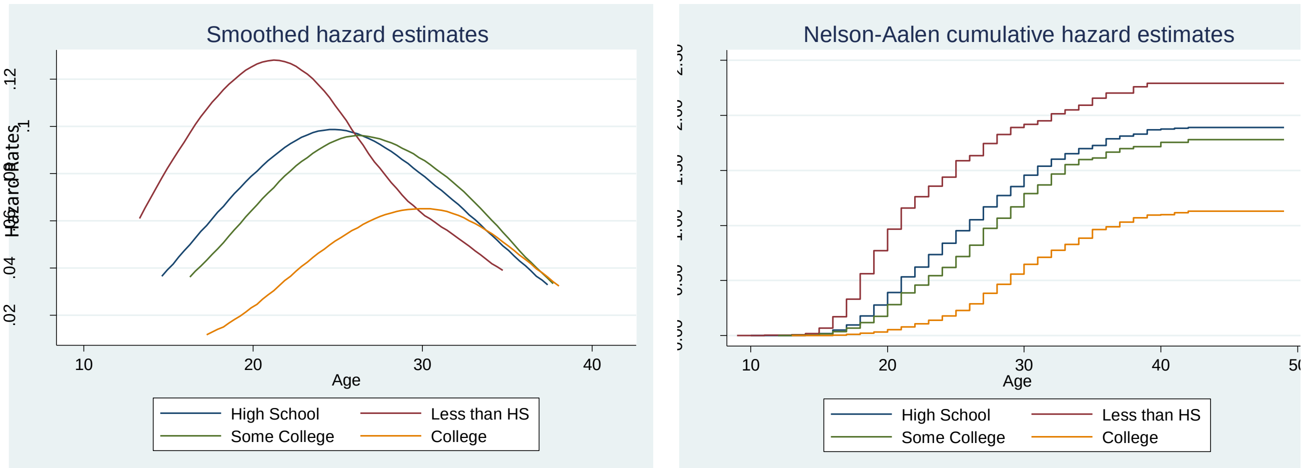
fecundability still at risk. Because the number of women having a first birth at older ages is relatively small, the probability calculations, and thus, the hazard estimates, are not always reliable at the higher tail (Allison, 2010).

Figure 4. Cumulative Hazard Function by Racial/Ethnic Group: Age at Time of First Birth



Since education is highly associated with fertility behaviors, I analyzed the survival and hazard functions by educational attainment. Figure 5 shows the hazard rates and the cumulative hazard estimates by educational attainment without control variables. The line at the top, with highest hazard rates corresponds to those with less than high school education. The two lines in the middle correspond to those with high school degrees and those with some college education respectively, and the bottom line corresponds to those with college degrees. At young ages we can see that women with lower education have a higher risk of having a first birth than those with higher education. The Log-rank and Wilcoxon tests confirm that these differences are highly statistically significant (p < .001).

Figure 5. Smoothed Hazard and Cumulative Hazard Function by Educational Attainment



The risk of having a first child peaks during the early 20s for those with the lowest education. In our sample, the median age at first birth for women with less than high school education is 18, followed by 21 for those who graduated from high school, 22 for those with some college education, and 27 for those with a college degree. The highest differences in the mean at age of first birth are at the two extremes: those with less than high school education have their first child three years younger on average compared to the next higher category (high school graduates); and those with college education have their first child four years older on average compared to the previous lower category (those with some college education). The difference between the two groups in the middle is only one year.

Cox Regression for Age at Time of First Birth: Analyzing Early Childbearing

As other scholars have noted, applying OLS regression to predict the age at time of first birth poses a series of problems. OLS regression cannot account for censored cases, and thus, childless women have to be removed from the analysis leaving us with a potential problem of sample bias. In addition, age at time of first birth is not normally distributed as far more women have their first child in their early 20s than in their late 30s. Survival models such as Cox regression represent an alternative methodology to overcome these problems (Allison, 2010; Raftery, et al., 1996). Since most of the data available in the GSS is in years, the use of discrete-time survival analysis is the most convenient way to analyze age at time of first birth. Table 2 shows the regression results predicting the age-specific hazard rate of having a first birth using a discrete time Cox model[3]. I used the *strata* option for educational attainment to allow the hazard to be different for each level of education and to interact with time. By stratifying the sample, the hazard estimates control for educational attainment, but without imposing the assumption that the effect of education is constant over time[4] (Allison, 2010, p.179-183; Cleves, Gould, & Gutierrez, 2010).

Table 2 reports the estimated coefficients as well as the hazard ratios or exponentiated coefficients. Our results indicate that race has a significant effect on the age-specific risk of having a first child, controlling for education and other factors. As we can see, most of the coefficients for race/ethnicity are significant, except for the coefficient for Asians. Controlling for education and other factors, Asians do not have a significantly different age-specific risk (or hazard) of having a first birth compared to Whites; all other minority groups have a significantly higher risk of having a first child. African Americans have a 45% greater age-specific risk (or hazard) of having a first birth than Whites, similarly, Mexicans have a 28% greater risk, whereas other Hispanic groups have a 60% greater risk than Whites of having a first child, holding constant other factors.

According to this model there are some small significant differences on the hazards by age cohorts controlling for other factors. Women who are between 25 and 29 years old have the same risk of having a first birth than those who are 20 to 24. However, the risk of having a first birth for women ages 30 to 34 who have not previously given birth is 31% higher than that of

[3] An exact marginal method for handling ties was used to calculate the coefficient estimates.
[4] The only drawback of this strategy is that the coefficients for the variable used to stratify to sample (education in this case) are not reported in the output.

women ages 20 to 24. Similarly women ages 35 to 39 who have not previously had a child have a 23% greater risk of having a first birth than woman ages 20 to 24. After age 40 the hazard rate is not significantly greater than that of women ages 20 to 24.

Table 2. Estimated Coefficients and Hazard Ratios Based on Cox Regression for Birth of First Child : GSS 1994-2008 (N = 7232)

Independent Variables	Coeff.		S.E.	z	Haz. Ratio	
Race/Ethnicity						
White (referent)	----		----	----	----	
Black	0.374	***	0.048	7.74	1.453	***
Mexican	0.245	**	0.080	3.07	1.278	**
Other Hispanic	0.470	***	0.092	5.12	1.600	***
Asian	0.043		0.092	0.47	1.044	
Other	0.267	***	0.074	3.62	1.307	***
Age group						
20 - 24 (referent)	----		----	----	----	
25 - 29	0.052		0.095	0.54	1.053	
30 - 34	0.272	**	0.091	2.98	1.313	**
35 - 39	0.209	*	0.091	2.30	1.232	*
40 - 44	0.173		0.091	1.90	1.189	
45 - 49	0.162		0.092	1.76	1.176	
Lived with both parents at 16	-0.092	*	0.036	-2.54	0.912	*
Number of siblings	0.020	***	0.005	4.20	1.020	***
Family Income[1]	0.028	***	0.006	4.83	1.029	***
Raised as Catholic	-0.134	***	0.038	-3.52	0.874	***
Survey year	-0.001		0.004	-0.38	0.999	
Log likelihood	-13356.51					
df	15					
BIC	26846.3					

* P < .05   ** P < .01   ***P < .001   Stratified by Education

[1] Income in $10,000 in constant dollars.

In this model I also control for nuclear family upbringing, that is, whether the respondent lived with both parents at age 16. Growing up in a nuclear or two-parent family reduces the hazard (risk) of having a first birth by 9%, ([.912-1]*100% = -8.8%) independent of the woman's age, race, education, and other variables. This may indicate a protective effect of nuclear family upbringing against early pregnancy. Family size has also been considered to have a significant effect on the timing of childbearing. In this model I found that number of siblings has a small but significant effect. Each additional sibling is associated with a 2% increase in the risk of having a first child, controlling for other factors. Family income is used as a control variable, and it seems to also have a small, but significant effect. I introduced income squared to test whether the effect of income was curvilinear, but the squared term was not significant. In general, for each $10,000 increase in real family income, the age-specific hazard of having a first child increases by 3%.

Previous studies also indicate that religious affiliation may have an effect on childbearing initiation. Using early waves of the GSS, Hout, Greeley and Wilde (2001) concluded that mainline Protestants adopted birth control early, whereas Catholics and conservative Protestants, such as Evangelical groups, were slow to adopt and even repealed birth control methods, a fact that explained early childbearing and higher fertility rates among the latter groups. Conversely, in this model we see that females who were raised as Catholic have a 13% lower age-specific risk of having a first child compared to those who were not raised as Catholic. However, it is important to note that the referent group includes not only Protestants, but also other religious groups such as Evangelicals, a denomination that predominates among the African American population, and that also has a significant representation among some Hispanic groups. However, as previously mentioned, using the GSS it was not possible to separate Evangelicals from other religious groups. Finally, I introduce the year of the survey to control for period effects; however this variable was not significant.

For the most part, this analysis shows that there are significant differences in the risk of having a first child across racial/ethnic groups. I conclude that race has a significant effect on the initiation of childbearing that is highly significant even after we control for education, income, family upbringing, and other variables. Most minority groups, namely, African Americans, Hispanics and other minority groups, have a significantly higher age-specific risk of having a first child compared to Whites, controlling for the effects of education and other variables. The model also shows that Asians, despite exhibiting a significantly higher median age at first birth, are not significantly different from Whites in their risk of having a first child after controlling for education and other variables.

After we control for education and other factors, we observe that White and Asian females tend to delay childbearing more than other ethnic groups. Immigration scholars have indicated that the selective nature of immigration can explain why Asians behave differently from other immigrant groups, as Asian immigrants tend to be better educated and disproportionally urbanites, both of which are characteristics associated with delays in childbearing and lower fertility. Other authors also affirm that the family formation behavior of certain Asian groups, such as Chinese women, is strongly influenced by the 'later-longer-fewer' campaign (later marriage, longer wait, and fewer children) and the 'one-child' policy that were put in place in China in 1976 and 1979, respectively, to promote late childbearing and lower fertility (Hwang & Saenz, 1997). Many scholars affirm that fertility policies in China have had a tremendous effect in the Chinese population and Chinese immigrant groups, including poor families as well as those with low levels of education (Kasinitz et al., 2008). Due to sample size limitations, I could not separate Chinese from other Asian groups, however this phenomenon may explain part of the different pattern we observe between Asians and other recent immigrant groups.

Other scholars have explained the early age of childbearing initiation among Hispanic groups by adducing cultural values brought from their country of origin such as 'familism' or values that encourage girls to form families at an early age and that place low value on education (Landale et al., 2009). However, this theory cannot explain why African Americans also exhibit similar high levels of early childbearing. Recent studies instead conclude that ethnic stratification shapes structural forces and opportunities that explain high rates of early childbearing among disadvantaged groups, including Hispanics and African Americans (Frank & Heuveline, 2005). Due to their disadvantaged economic situation, Hispanic and African American children are encouraged to start working at early ages, preventing them from achieving higher levels of education. This situation also affects children's expectations about their future, which is seen as not incompatible with starting families at an early age (Zhou, Lee, Agius Vallejo, Tofoya-Estrada, & Xiong, 2008). Educational expectations play an important role as those who expect to attend college are more likely to delay childbearing (Landale et al., 2009).

In addition early childbearing among Hispanic groups have also been explained by certain religious norms that support abstinence and censure the use of birth control methods, such as condoms or contraceptives, failing to prevent early childbearing among females with these religious beliefs. Some scholars also found that information about sex is limited among Catholic and Evangelical families (Wildsmith & Raley, 2006; Kasinitz et al., 2008). Conversely, Asian origin groups generally do not have religious opposition towards contraceptive methods or abortion. Moreover, while abortion is illegal in most Latin American and Caribbean countries, it is legal and widely available in China and other Asian countries. These explanations can also account for some of the differences in early childbearing between Hispanic and African Americans on the one hand, and White and Asian groups on the other hand (Kasinitz et al., 2008).

Fertility Differentials across Racial/Ethnic groups

To analyze age-specific fertility, I used total number of children ever born. This variable does not follow a normal distribution, but a Poisson distribution for small counts. Thus, I used a Poisson regression instead of an OLS regression. A Poisson model ensures that the expected count (number of children) will always be nonnegative, and it also solves the problem of assuming constant variance of the errors (homoscedasticity) in OLS regression (Long, 1997, p. 223). Because many of the birth histories are censored at the time of survey, I control for women's time exposure to fertility using age as the time of exposure. The exponential feature of the exposure option in Poisson regression also allows for a nonlinear effect of age on fertility, letting the effect of age to vary by different levels of expected number of children. The dependant variable is the number of children ever born, and the time of exposure is the age at the time of survey[5].

Table 3 presents the results[6]. Poisson regression predicts the rate of occurrence of the event (having children) per unit of exposure time (Long, 1997). Table 3 displays incidence rate ratios (IRR) or exponentiated coefficients. The results are pretty similar to what we observed in the analysis for having a first child. Race/ethnicity has a strong significant effect on the predicted number of children. Model 1 shows that African Americans have on average nearly 60% more children than Whites, Mexicans have 70% more children, and other Hispanic groups have 68% more children than Whites controlling for cohort differences. However, as we will see later, part of this race effect is spurious, and can be explained by education differences. Asians again were not significantly different than Whites in their fertility rates in any of the models.

All of the age cohort differences were significant. Women who at the time of the survey were 25 to 29 years old had on average 67% more children than those who were 20 to 24 years

[5] I also tried starting the time of exposure at age 12, and the results were very similar to the ones shown in Table 3.
[6] I tested the models for over-dispersion using a negative binomial regression, and the parameter alpha (that measures over-dispersion) was not significant, suggesting no presence of over-dispersion in the models.

Table 3. Incidence Rate Ratios (IRR) Estimates Based on Poisson Models: GSS 1994-2008. DV = Number of Children ever Born (1-3). DV = No. of Children Beyond the 1st (4)

	Independent Models (N = 7232)						Conditional Model (N= 5156)	
Independent Variables	1		2		3		4	
Race/Ethnicity								
White (referent)	----		----		----		----	
Black	1.595	***	1.379	***	1.397	***	1.105	**
Mexican	1.705	***	1.339	***	1.334	***	1.160	**
Other Hispanic	1.683	***	1.413	***	1.523	***	1.137	*
Asian	0.945		1.065		1.031		1.153	
Other	1.603	***	1.372	***	1.299	***	1.084	
Age group								
20 - 24 (referent)	----		----		----		----	
25 - 29	1.669	***	1.860	***	1.733	***	1.084	
30 - 34	2.423	***	2.727	***	2.459	***	0.941	
35 - 39	2.499	***	2.856	***	2.578	***	0.740	***
40 - 44	2.553	***	2.875	***	2.604	***	0.565	***
45 - 49	2.477	***	2.838	***	2.584	***	0.441	***
Education								
Less than High School			1.514	***	1.433	***	1.147	***
High School (referent)			----		----		----	
Some College			0.874	**	0.901	**	0.994	
College			0.509	***	0.542	***	1.044	
Lived with both parents at age 16					0.933	***	1.058	*
Number of siblings					1.024	***	1.014	***
Family Income[1]					0.994		1.007	
Married					1.433	***	1.157	***
Employed					0.795	***	0.753	***
Raised as Catholic					0.899	***	1.036	
Survey year	1.004		1.006	**	1.004		1.011	***
Age (exposure)							Age - Age at 1st birth (exposure)	
Log likelihood	-13319.0		-12719.8		-12421.0		-6647.9	
df	12		15		21		21	
BIC	26744.6		25572.9		25028.6		13475.4	

* P < .05 ** P < .01 ***P < .001

[1] Income in $10,000 in constant dollars.

old controlling for race. All the other older cohorts have nearly 2.5 as many children as do females age 20 to 24. These results may reflect the fact that older women have had more time to bear children than the younger cohort.

Model 2 controls for education and shows a significant decrease in the coefficients and the IRRs for race. More than a third of the race effect on fertility can be explained by differences in education. As seen in Model 2, controlling for education and age cohort, African Americans now have 38% (compared to 60% in Model 1) more children than Whites. Mexicans have 34% (compared to 70% in Model 1) more children, and Other Hispanic have 41% (compared to 68% in Model 1) more children than Whites.

The changes in the age cohort coefficients are smaller. The coefficients for education show that women with less than high school education have on average 51% more children than those who graduated from high school, controlling for race, cohort, and period effects. Those with some college education have 13% fewer children ([.87-1]*100% = -13%) than those with high school education, whereas those with college degrees have about half the number of children than those with high school education do. Again, since we lack information on the timing of education, we cannot affirm whether education affects childbearing decisions or whether childbearing affects educational outcomes. However, although the directionality of this effect can be debatable, the associations are strong and highly significant.

In Model 3, I include other control variables, however the coefficients for race show only minor changes. Race has still a strong significant effect on fertility. As previous studies have also found, most racial minority females exhibit higher fertility rates than Whites (Landale, et al., 2010; Frank & Heuveline, 2005). Some scholars have adduced cultural pro-natalist values to explain higher fertility rates among the Mexican origin population. Other scholars have argued that there is a 'selection effect' that occurs prior migration by which Hispanic women who decide to migrate already present specific socio-demographic profiles that contribute to explain their higher fertility (Frank & Heuveline, 2005; Feliciano, 2005).

Model 3 also shows the effect of nuclear family upbringing and other variables on fertility. Women who were raised in a nuclear family have on average 7% fewer children than those who were not raised by both parents, independent of the effects of other factors. Each additional sibling is associated with a 2% increase in the predicted number of children. Family income does not have a significant effect on expected number of children in this model. Marital status also has a significant positive effect on fertility. The coefficient indicate that those who are married have on average 43% more children than those who are single, divorced, separated or widowed. Since the GSS does not provide data on age at marriage anymore, we do not know whether women married before or after they had their children, however there is a strong positive association between marital status and fertility. It is possible that marriage encourages women to have more children, or that women who have more children seek to get married, or maybe we have a combination of both scenarios. In the recent decades it has become crucial to analyze the effect of cohabitation on fertility. The GSS does not provide information on cohabitation status and it is not clear whether cohabitors are included in the single or married category. Thus, a study with recent data on cohabitation becomes necessary to complement these results.

Employment status is negatively associated with fertility. Women who are employed have on average 20% fewer children than women who do not work. Due to data limitation we cannot analyze the directionality of this effect. It is possible that women who have more children are impeded to participate in the labor force, but it is also very likely that women who are employed (or expect to be employed) limit the number of children they have in order to free time and be able to work. There also seems to be an effect of being raised as Catholic. Women raised as Catholic have on average 10% fewer children than those who were not raised as

Catholic, however, as mentioned before, the referent group could be conflating Evangelicals with Protestants, and this study does not separate these groups. This effect could also be the result of comparing Catholics to other groups including other religion and non religious groups, that combined represent about 14% of the sample. A more detailed analysis of the effect of religion becomes necessary to disentangle these possible effects.

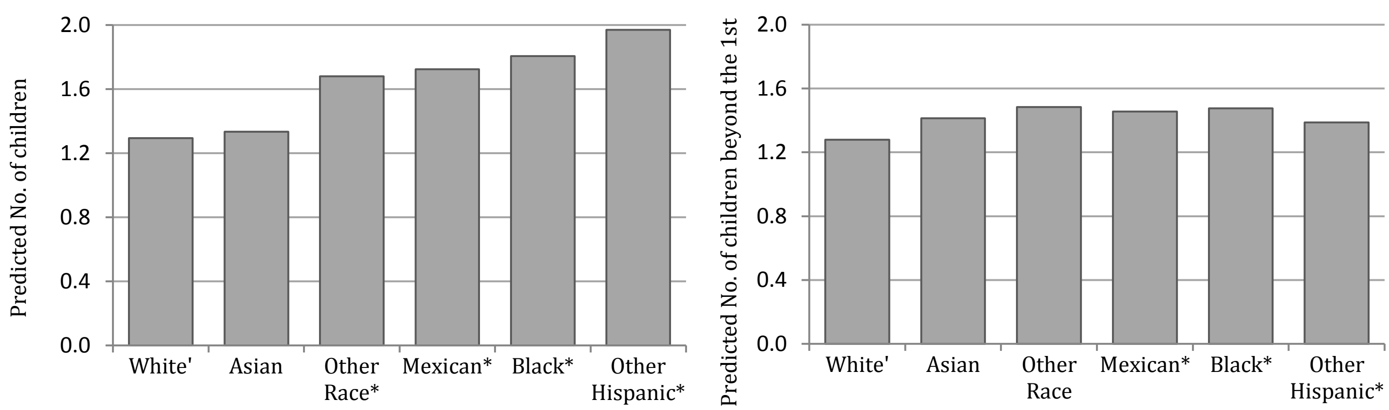
In order to further analyze the effect of age at time of first birth on total number of children, I ran a conditional Poisson model for females who have had at least one child (Model 4). Women who do not have children are omitted from this model. The dependant variable is the number of children beyond the first one, and the exposure time is the years between the age at time of first birth and the age at the time of the survey. Thus, this model controls for age at time of first birth for women who have had a first child. The results are shown in Model 4. As this model shows, the race effects on fertility are greatly reduced from those shown in Model 3. According to the conditional model, African Americans only have 11% more children beyond the first one than Whites (compared to 40% as seen in Model 3), controlling for cohort effects, education, marital status, and other factors. Mexicans have on average only 16% more children than Whites controlling for other variables. Other Hispanics have 14% more children than Whites (compared to 52% more in Model 3). Asians are not significantly different than Whites in the number of children they have beyond the first child, controlling for other factors.

Figure 6 compares the expected number of children by race/ethnicity based on the coefficients of the Poisson regression in Models 3 and 4, and setting all other variables to their mean. The calculations are based on age-specific fertility, not completed fertility. Panel (a) includes childless women, whereas Panel (b) does not. We can observe greater differences in age-specific fertility across ethnic groups compared to Whites in panel (a), however, once we account for the age when women start bearing children in panel (b) the differences are smaller and not all of them are significant. This finding suggests that cultural explanations are more suitable to explain racial differences in the timing of childbearing initiation, but culture does not have a strong explanatory power to explain age-specific fertility differentials across racial groups once age at first birth is taken into account.

Figure 6. Predicted Number of Children based on Poisson Regression by Race/Ethnicity Adjusting for Other Factors[1]

(a) Predicted number of children (Model 3)

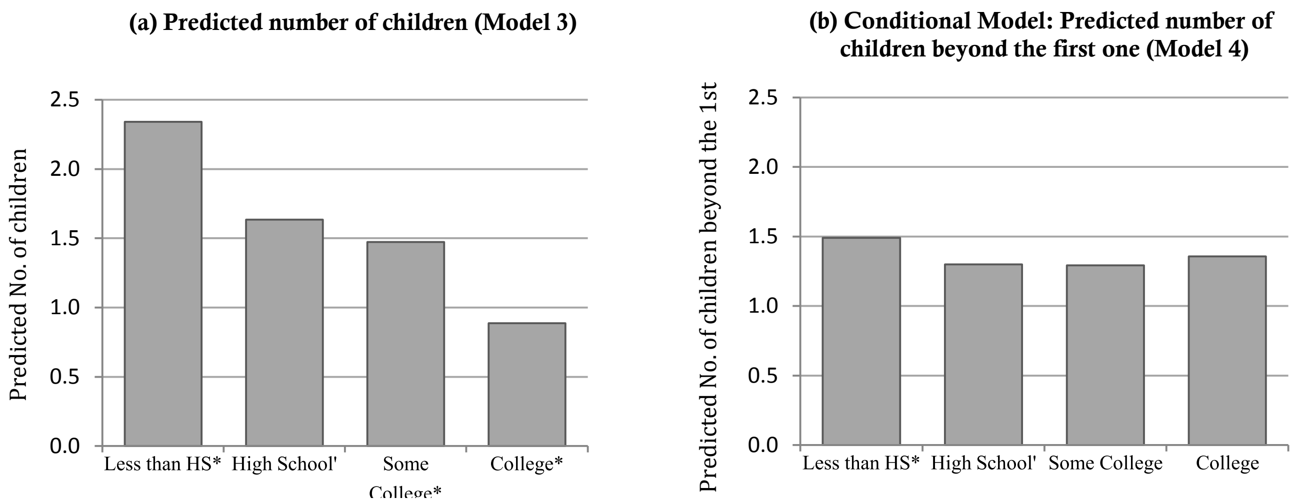
(b) Conditional Model: Predicted number of children beyond the first one (Model 4)



[1] Setting all other variables in the model to their mean.
* Statistically significant differences.
' Referent category.

In Model 4, cohort differences are significant only for the three older cohorts, however the effects are reversed. Women who are 35 years or older have on average fewer children beyond the first one compared to 20 to 24 years old women who have already had a child. These results are in accordance to what other scholars have noted that higher order births become less frequent as women age (Preston et al, 2001). It is very interesting to note that the coefficients for education either disappear or are greatly reduced in this conditional model. Holding constant other factors, women with less than high school education have only 15% (compared to 43% in Model 3) more children beyond the first one than women who graduated from high school, and this effect is statistically significant ($p < .001$). However, the coefficients for those with some college education and those with a college degree totally disappear and become insignificant. Figure 7 shows the predicted number of children by educational attainment based on Models 3 and 4. These results suggest that if there is an effect of education on fertility, it is stronger and significant to explain childbearing initiation, but this effect is no longer strong when we want to explain total number of children beyond the first child controlling for the mother's age at first birth.

Figure 7. Predicted Number of Children based on Poisson Regression by Educational Attainment Adjusting for Other Factors[1]



[1] Setting all other variables in the model to their mean.
* Statistically significant differences.
' Referent category.

In Model 4 the effect of nuclear family is reversed and its significance is reduced. Females who were raised in a nuclear family have on average 6% more children beyond the first one than those who did not grow up with both parents, controlling for other factors. The effect of number of siblings on age-specific fertility is also greatly reduced; each additional sibling is associated with having 1.4% more children beyond the first one, controlling for other factors. Family income remains not significant. The coefficient for being married is also diminished. Those who are married have on average only 16% (compared to 43% in Model 3) more children beyond the first one than women who are not married, controlling for other factors. The effect of being employed remains mainly unchanged and highly significant. The effect of Catholic upbringing disappears and becomes non significant. As mentioned before all these models include the year of survey to control for possible period effects; however, this variable is either not significant, or has a small effect.

In sum, these analyses reveal that most of the race differentials on fertility behaviors are explained by the women's age at the time of first birth, which seems to affect the total number of children women bear. When we control for age at first birth in the conditional model, a significant part of the race effect on fertility disappears. Thus, I conclude that race differences in fertility behaviors are due, in great part, to the timing of motherhood initiation, and that the racial/ethnic differences in the total number of children is small after we control for women's age at time of first birth. Similarly, there is a strong association between education and fertility, but this relationship weakens when we control for the age of women at first birth. Finally, the effect of being raised in a nuclear family seems is strong and significant, and it seems to prevent early childbearing, and reduce fertility, but after taking into account mothers' age at first birth, being raised by both parents is instead associated with having more children.

Next Steps

Consistent with what previous studies have found, my analysis confirms that there is a strong effect of race/ethnicity on fertility behaviors. The preliminary assessment using survival analysis indicates that, on average, African American and Hispanic women bear their first child at younger ages compared to White women. Using a discrete time Cox model, I found that African American, Hispanic, and other racial minority groups have a significantly greater risk of having a first child compared to White women, whereas Asians, although they exhibit a higher median age at first birth, are not significantly different from Whites when education and other variables are taken into account. Similarly, most ethnic minorities also exhibit higher age-specific fertility compared to Whites. A more detailed analysis using a conditional Poisson regression reveals that a great part of the effect of race on fertility is explained by the age when women have their first child. However, this analysis will be reproduced disaggregating Asians, Mexicans, and other Hispanics into native born and immigrants. In addition, region and the population size where respondents were interviewed will be included in order to control for urban contexts.

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