# Premarital Birth among Young Hispanic Women: Evidence from Semiparametric Competing Risks Analysis

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## Abstract

The existing literature on welfare effects on marriage and fertility has largely focused on groups of white and black women. By contrast, the group of Hispanic women has received little attention. This paper examines the effects of welfare generosity on young woman's premarital fertility and marriage choices for a sample of young Hispanic women. A bivariate competing risks duration model framework allows us to identify the process of young women's premarital fertility and the process of marriage, effectively controlling for observed characteristics and unobservables. Our findings indicate that a 10% increase in welfare generosity would have resulted in a 10% increase in premarital births and a 7% decrease in marriages by age 24; both effects are significant.

*Classification*: I00, J10.

*Keywords*: Welfare generosity; Premarital fertility; First marriage; Unobserved heterogeneity.

## Introduction

The Aid to Families with Dependent Children (AFDC) program, which was replaced by Temporary Assistance for Needy Families (TANF), was often blamed for increasing illegitimate births, declining marriage, and rising divorce. The main argument was that, by favoring single-parent families, the AFDC welfare system reduced the costs of having children and increases the value of single parenthood as an alternative to marriage.

Between 1990 and 1999, the proportion of welfare families that were minorities increased from 60 percent to just over 67 percent and the Hispanic welfare population grew from 17 percent to 25 percent (U.S. Department of Health and Human Services, 2005). However, the existing literature on the effects of welfare on marriage and fertility has largely focused on groups of white and black women. In a review of 112 studies on the subject, Moffitt (1998) refers to the word “Hispanic" only twice in relation to the papers by Plotnick (1990) and Lichter *et al*. (1997). Plotnick (1990) uses NLSY data to analyze the link between welfare and teenage out-of-wedlock childbearing using a simple logit hazard framework. He does not find a strong link. His analysis does not allow for state fixed effects or unobserved heterogeneity and marriage is treated as random censoring. Schultz (1994) uses data from the 1980 U.S. Census to analyze probit equations on the effect of welfare on the Probability of a Woman Living with a Spouse and Tobit and OLS equations to analyze the effects of welfare on the number of children ever born. He considers blacks and whites and three age groups. He finds that AFDC generosity has a negative, statistically significant, effect on marriage for the 14 to 24 age group. The effect of AFDC on fertility is negative and statistically significant only for whites ages 15 to 24. He does not consider state fixed effects. Hoynes (1997) uses data from the Panel Study of Income Dynamics (PSID) to also study the effect of welfare generosity on the probability of female headship. She finds no significant effect of welfare for white women; for black women the effect became insignificant after controlling for individual unobserved heterogeneity. Lichter *et al.* (1997) consider the impact of welfare on the percentage of female-headed families. They find that welfare had only a small impact. Their analysis blends different sources of female headship (like premarital birth and divorce) and considers race dummies as explanatory variables. Rosenzweig (1999) uses NLSY data to study the effect of welfare on the probability that a woman has a premarital birth versus only marital births or no births by age 22. After controlling for state fixed effects and cohort fixed effects, he finds a significant and quantitatively large positive effect of welfare on the probability of premarital fertility, especially among low income women. Because his analysis is only concerned with the status quo at a specific age, it cannot distinguish between the hypothesis that the increase in premarital birth is the result of a direct effect of welfare on teen fertility, and the hypothesis that welfare reduces the likelihood of marriage. Any one of these effects can increase premarital births. In a recent paper, Keane and Wolpin (2007) structurally estimate a dynamic programming model of life-cycle decisions of young black, Hispanic and white women. They account for a number of labor and marriage market outcomes, but do not focus on premarital birth. While TANF has substituted AFDC, the relative simplicity of the old welfare program provides a unique opportunity to analyze the effects of economic incentives on marriage and fertility behavior.[1]

In this paper, we analyze the effects of the AFDC welfare generosity on a young woman's premarital fertility and marriage choices for a sample of young Hispanic women from the National Longitudinal Survey of Youth (NLSY). A bivariate competing risks duration model framework allows us to identify the process of young women's premarital fertility and the process of marriage. To evaluate the sensitivity of the empirical results, we estimate a variety of econometric specifications for different sets of explanatory variables, state fixed effects, and individual specific unobserved heterogeneity. We also provide a comparative analysis using a sample of white and black young women. Overall, our results for the black and white samples are consistent with previous research and we interpret this as validation of our

[1] Blank (2002) provides a review of more recent literature, including studies examining the impact of welfare reform on female headship.

empirical strategy.

The empirical evidence indicates that welfare has a significant effect on a woman's decision to marry and a woman's premarital birth decisions. Results from a policy simulation indicates that a 10% increase in welfare generosity would have resulted in a 10% increase in premarital births and a 7% decrease in marriage by age 24; both effects are significant. Our analysis also suggest that a significant proportion of early marriage and premarital births for whites and blacks is due to individual specific characteristics unobserved to the econometrician, or unobserved heterogeneity, while this is not the case for Hispanics.

The next section describes the data. This is followed by a brief description of the analytical framework and the econometric model. After presenting the main estimation results for the Hispanic sample, including simulation results that illustrate the quantitative impact of welfare, we then discuss the racial differences that emerge from a comparative analysis. Finally, concluding remarks are provided.

## The Data

Teenagers and older women may have different reasons to conceive out of wedlock. In addition, we want to focus our attention on the problem of premarital birth among teenage women. The evidence suggests that this group of women bears the highest burden as a result of a premarital birth (Hotz *et al.,* 1997; Ribar, 1996; Haveman *et al*., 1997). We consider a sample of 886 Hispanic women from the main sample and the supplemental sample of Hispanic women in the NLSY. Individuals in the sample have been interviewed every year since 1979 and extensive information on family background, fertility and marital history has been collected. In addition, retrospective fertility and marital histories were collected in 1979. We follow women until the time of their first transition into marriage or premarital birth or until age 24 whichever comes first.

A premarital birth spell is defined from the time the young woman becomes at risk of having a child, at age thirteen,[2] until the time of the first premarital birth. Similarly, the first marriage duration begins at age thirteen and continues until the time of the first marriage. In the empirical analysis, we classify births from cohabitating relationships as premarital births. Unfortunately, the information available in the NLSY about cohabitation is not as detailed as the information on marriage and fertility. At the time of each interview, the woman is asked if she is cohabitating with a partner of the opposite sex. There is no information about the duration of the cohabitation and there is no retrospective information on cohabitation for the years prior to 1979. We have examined data on the youngest cohorts in the NLSY to study the extent of cohabitation. Overall, cohabitation appears to be a long-term alternative to marriage for only a small percentage of the women in our sample who experience a premarital birth.[3] Thus, given data limitations we feel that not accounting for cohabitation in our empirical work is a reasonable compromise.

For each individual, the NLSY provides the state of residence at age 14 and the state of residence for each survey year. However, the state of residence between age 14 and the age at the time of their first interview is unknown. Following Rosenzweig (1999), we substitute the missing information with the state of residence at age 14, or at the age of the first interview depending on which one is closest in time. We will use this information later on to associate state dummies to each individual-year observation. Another shortcoming of the NLSY data set is that it does not provide information on parental income. As in Rosenzweig (1999), a measure of parental income is constructed by combining information on the characteristics of the parents of the women in the sample when they were age 14 with information on

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[2] No women in our sample had a child before age thirteen.

[3] Moffitt *et al*. (1998) analyze the extent of cohabitation among welfare recipients. They find that a significant proportion of single mothers cohabitate. Our results are consistent with their findings since we focus our attention on young women at the time of their first premarital birth while their sample includes all single women.

median wages by occupation, education level, gender, race and marital status, obtained from the census. Finally, the NLSY data is supplemented with data on welfare generosity, the availability of abortion services, and wages at the state level.

Table 1 reports descriptive statistics for the relevant variables in our Hispanic sample. Some of the variables describe characteristics of the household in which each woman in the sample resided at age 14. About seventeen percent of Hispanic women were living with a single mother and about sixty-eight percent of Hispanic women were living with both parents at age 14. Table 1 also provides descriptive statistics for other important variables that will be used in the estimation of the model, like the standardize Armed Forces Qualification Test (AFQT),[4] religiosity and urban residence. There are no important observable differences between women that experience a premarital birth and women that experience a marriage first. The second group of women live in households with slightly lower average income and education, and are less likely to live in an urban area. These women also have a lower standardized AFQT, are more likely to live with both parents and are less likely to live with a single mother.

[Table 1 about here]

AFDC payments have been declining consistently over the time period of analysis. Older cohorts in the NLSY have been exposed to a more generous welfare program while there is a higher proportion of premarital births among younger cohorts (Rosenzweig, 1999). We have also examined a composite measure of welfare generosity defined and used by Moffitt (1994). Because participants on the AFDC program usually participate in other welfare programs, Moffitt's measure of welfare generosity attempts to capture the overall value of welfare by including AFDC, Food Stamps and Medicaid Benefits. Moffitt's

[4] The test measures arithmetic reasoning, word knowledge, paragraph composition, and numeric operations. It was administered to most of the subjects in the NLSY and is commonly used as a measure of unobserved ability.

measure of welfare generosity for a family of two has remained relatively constant over the period of interest. We will use Moffitt's measure of welfare generosity in our analysis.[5]

Table 2 reports descriptive statistics of the process of marriage and premarital birth for three ethnic groups: Hispanic, Black and White. The average ages of premarital birth and first marriage are very similar for Hispanic and Black women and is somewhat higher for White women. However, there are important differences in the proportion of women in the marriage and premarital-birth states for different age groups. The process of marriage among Hispanic and White women seem to be very similar until age 22 but by age 24 we observe a 7% higher proportion of marriage among white women; black women are less likely to marry. In contrast, the proportion of Hispanic women that have experienced a premarital birth by a certain age is significantly lower than the proportion of Black women, and significantly higher than the proportion of White women. Thus, we believe that the process of marriage/premarital-birth among Hispanic women is sufficiently different from that of other ethnic groups to require separate study.

[Table 2 about here]

## Model

### *Analytical Framework*

To elaborate on a theoretical framework, consider a simple model illustrative of how single, childless, young women can become single mothers or married women. Assume that there are two types of men: marriageable men and sperm donors. Marriageable men are those that are acceptable as husbands, sperm donors are not. Define $M$ as the percentage of marriageable men and $1-M$ as the percentage of sperm donors. Define also $m$ as the marriage event and $\overline{m}$ as the sperm donor event. Assume that women value marriage but value children as well, even out of wedlock. Women search for marriageable

[5] This measure of welfare is defined in Moffitt (1994) page 626, footnote 12. Like Moffitt we did not find significant differences when using the AFDC variable instead of Moffitt's index of welfare generosity.

men with intensity $\eta_m$ and for sperm donors with intensity $\eta_{\bar{m}}$. For example, women that do not consider having children out of wedlock will have an associated $\eta_{\bar{m}}$ of zero, and women that are not interested on marriage will have an associated $\eta_m$ of zero. Consistent with search theory, the hazard of marriage, $h_m$, is equal to $\eta_m M$, or the product of the arrival rate of mates multiplied by the probability that this mate be acceptable as a husband. Similarly, the hazard of premarital birth is $\eta_{\bar{m}}(1-M)$.

The effect of welfare on increasing the value of single motherhood is well understood. However, the different ways in which welfare can affect marriage are perhaps less apparent. By increasing the value of single motherhood welfare increases the minimum quality required from a marriageable man and, therefore, results in a decrease in the percentage, $M$, of marriageable men. This is what we call the *direct effect*. This decrease in $M$ implies a decline in the likelihood of successful search for a marriageable man and, as predicted by the search theory, a decrease in the intensity of search for marriageable men, which in our model implies a decrease in $\eta_m$. This is what we call the *indirect effect*. In particular, if the cost of search is higher than its expected return a rational agent will stop searching altogether. Both the direct and indirect effects result in a decline in the hazard of marriage.

The empirical framework considered in this paper can be interpreted as a reduced form of a model of marital search in which women search for marriageable men and at each point have to decide whether to conceive a child out of wedlock, or to postpone fertility.

*Empirical Model*

We follow a young woman's marriage/premarital-birth process from age 13 to age 24 or until the time of censoring, whichever occurs first. We analyze this process within the framework of a competing risk model. This allows us to distinguish between the process of young women's premarital fertility and the process of marriage.[6] Some covariates, like welfare benefits or wages, change values from year to year in our data. The estimation approach allows for time-varying covariates with possibly varying associated parameter coefficients. Let $j=1$ denote first marriage and $j=2$ denote premarital birth. In the empirical specification the instantaneous probability, or hazard, of event $j$ at time $t$ is characterized by

$$\lambda_j(t_{ji} \mid x_{ji};\ \beta_j)\nu_{ji} = \lambda_{0j}(t_{ji})\exp\left(x_{ji}\beta_j\right)\nu_{ji}, \quad i=1 \ to\ N; \quad j=1,\ 2; \tag{1}$$

where $\lambda_{0j}(.)$ is the baseline hazard, $t_{1i}$ denotes the time (in months) to first marriage for the $i$*-th* young woman, $t_{2i}$ is the time to first premarital birth, and $x_{ji}$ represents a vector of explanatory variables (including time-varying covariates) associated with the $j$*-th* risk, with $\beta_j$ being the associated parameter vector. Here $\nu_i = \left(\nu_{1i}, \nu_{2i}\right)'$ denotes a vector of unobserved heterogeneity components that differ across risks. We observe data of the form $\left\{x_i,\ t_i,\ d_{1i},\ d_{2i},\ d_{ci}\right\}_{i=1}^{N}$, where $x_i$ represents a vector of explanatory variables, $t_i = \min(t_{1i},\ t_{2i},\ t_{ci})$, with $t_{ci}$ representing time of censoring and $d_{ji} = I\left(t_{ji} = t_i\right), \quad j=1,\ 2,\ c,$ a dummy variable. It remains to specify the joint density of unobserved heterogeneity components and the baseline hazard.[7]

Following Canals-Cerdá and Gurmu (2007), we employ a semiparametric estimation approach based on Laguerre series expansion of the joint density of the unobserved heterogeneity components.

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[6] This empirical framework can be interpreted as a reduced form of a model of marital search in which women search for marriageable men and at each point have to decide whether to conceive a child out of wedlock, or to postpone fertility.

[7] Cameron and Trivedi (2005, chapters 17-19) provide a good textbook treatment of hazard rate or survival models, including competing risks and economic applications.

Dependence between time till first marriage and time to first premarital birth can be introduced by allowing for dependence between $(\nu_{1i}, \nu_{2i})$. Following their approach, the unobserved heterogeneity components are approximated by a bivariate density of the form

$$g_N(\nu_1,\ \nu_2) = g(\nu_1)g(\nu_2)\frac{1}{\phi_N}\left[\sum_{k=0}^{K}\sum_{r=0}^{K}\rho_{kr}P_k(\nu_1)P_r(\nu_2)\right]^2, \tag{2}$$

where, suppressing reference to the $i$-*th* observation, $g(\nu_1)$ and $g(\nu_2)$ are gamma baseline densities with parameters $\alpha_1$ and $\alpha_2$ of $\nu_1$ and $\nu_2$, respectively, $\{P_k(\nu_1), P_r(\nu_2)\}$ are the associated Laguerre polynomials, $\phi_N$ is the constant of proportionality that makes $g_N(\nu_1,\nu_2)$ integrate to one, and $K$ denotes the degree of the polynomial. The parameter $\rho_{kr}$ may be considered as the coefficient of correlation between the $(k,r)$ polynomials in the bivariate distribution $g_N(\nu_1,\nu_2)$. After integrating out the effect of unobserved heterogeneity the ensuing mixture survivor function can be denoted as $S(t_{1i}, t_{2i} \mid x_i; \beta)$. The probabilities associated with first marriage and premarital birth can be expressed as

$$q_{ji}(t_i;\beta) = \frac{-\partial S(t_{1i},\ t_{2i} \mid x_i;\ \beta)}{\partial t_j}\Big|_{t_{1i}=t_{2i}=t_i}, \quad j=1,\ 2. \tag{3}$$

Likewise, for $j=c$, the probability of censoring is $q_{ci}(t_i;\beta) = S(t_{1i}, t_{2i} \mid x_i;\beta)$. A convenient feature of this approach is that the survival function and the probabilities that derive from it accept a close form analytical representation.

The baseline hazard is defined as a step function (Meyer, 1990). More precisely, we first divide duration time into $R$ intervals with endpoints $a_1,\ a_2, ...,\ a_R$, and then define $\lambda_{0j}(t_j) = \exp(\delta_{rj}) > 0$, for $a_{r-1} \le t_j < a_r$. Now, for $j = 1,\ 2,\ c$, let $\hat{q}_{ji}(t_i;\beta)$ denote the resulting probabilities obtained from this approach. The log-likelihood function can be expressed as

$$\mathbf{L}(\beta) = \sum_{i=1}^{N} \left[ d_{1i} \log \hat{q}_{1i}\left(t_i;\beta\right) + d_{2i} \log \hat{q}_{2i}\left(t_i;\beta\right) + d_{ci} \log \hat{q}_{ci}\left(t_i;\beta\right) \right], \quad (4)$$

where the vector $\beta$ consists of parameters associated with the baseline hazard, covariates, and distributions of unobserved heterogeneity.

The parameters of non-linear econometric models are generally difficult to interpret. Thus, we consider a policy simulation to estimate the average effect of welfare on premarital birth and marriage. Based on equation (3), we can specify

$$Q_m\left(t \mid x;\beta\right) = \int_0^t q_1\left(y \mid x;\beta\right) dy \text{ and } Q_f\left(t \mid x;\beta\right) = \int_0^t q_2\left(y \mid x;\beta\right) dy \quad (5)$$

as the probability of marriage before a certain age $t$ and before premarital birth, and as the probability of premarital birth before time $t$, respectively. This integral is computed numerically using Gaussian Quadrature techniques. Define

$$E_i\left(Q_j\left(t \mid x_i;\beta\right)\right), \quad \text{for } j = m, f,$$

as the average probability of marriage/premarital-birth over the whole population.

## Results

We consider twenty five different model specifications, with models differing in three possible dimensions. On the one hand, we consider two different sets of explanatory variables: what we call the small set consists of the variables welfare generosity, standardized AFQT test, mother's education, number of siblings, and family income, as well as two dummy variables, one associated with observations with missing mother's education and a second one that accounts for missing income information; what we call the large set consists of all the variables in a web table (Table A2) as well as three dummy variables, two associated with observations with missing mother's and father's education and a third one that accounts for missing income information.[8] On the other hand, we consider five different specifications for the unobserved heterogeneity component. Model 1 does not control for unobserved heterogeneity, in model 2 the series expansion is truncated at $K = 1,$ which corresponds to the case of a bivariate gamma distribution, in models 3 to 5 we consider series expansions truncated at $K = 2,\ 3$ and $4$, respectively. Finally, we also estimate the models with and without the inclusion of state fixed effects. In order to guarantee the consistency of our estimates, models that control for state fixed-effects are estimated using a smaller sample of individuals consisting of only those states with at least twenty five individuals in our sample in year 1979. Because the Hispanic population is highly concentrated, this reduces the total number of states to eight and the sample size to 725 observations.

*Main Findings*

Estimates of the parameters associated with the welfare variable and the hazards of marriage and premarital birth for young Hispanic women are reported in Table 3.[9] Economic theory suggests that welfare generosity increases the likelihood of premarital birth, and decrease the likelihood of marriage. As a result, we would expect welfare to have a positive effect on the hazard of premarital birth and a negative effect on the hazard of marriage.

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[8] We find that the large set of variables explain more than the small set. Taking into account that competing risk models are highly nonlinear, one possible explanation is that the large set of explanatory variables explains away some confounding effects in the marriage and fertility processes and this allows for a more precise estimate of welfare effects.

[9] Detailed results for explanatory variables other than welfare generosity are given in a web table, Table A1. Web tables A1 through A5 are available from the authors or at http://www2.gsu.edu/~ecosgg/research/pdf/cg_aej.pdf

[Table 3 about here]

As reported in Table 3, our results are consistent with the predictions of the theory in all model specifications considered. However, the coefficients associated with the welfare variable are not always significant. In particular, they are not significant, at the usual level, for all the models that include a small set of explanatory variables. Furthermore, when we add state fixed effects to these models the magnitude of the associated *t*-values decreases, although the magnitude of the coefficients increases. In contrast, for model specifications including a large set of variables the effect of welfare on the hazard of premarital birth is always significant, and the effect of welfare on the hazard of marriage is significant in the model without unobserved heterogeneity and the model with gamma unobserved heterogeneity, but becomes insignificant in models with more flexible forms of unobserved heterogeneity $(K > 1)$ although the magnitude of the coefficients does not change by much across models. Finally, after adding state fixed effects to these models the magnitude of the associated *t*-values decreases and the coefficients became insignificant, although the magnitude of the coefficients increase for the hazard of premarital birth, by as much as 50%, and decreases for the hazard of marriage.

Why do the effects of welfare became insignificant after the inclusion of state fixed effects? The introduction of fixed effects is presented by its proponents as a way to deal with the potential endogeneity of the welfare policy. However, a consequence of the introduction of state fixed effects is a reduction of the identifying variability in the welfare variable in several ways. On the one hand, fixed effects remove the variability across states, which leaves us with the over-time variability on welfare across individuals as the only source of identifying information. Notice that all the models estimated also include cohort specific dummies. In addition, the models that include fixed effects are estimated using a smaller subsample, since we restrict the sample to those states with at least 25 individuals in 1979. This results in a reduction in the variability of welfare across individuals as well as a reduction in sample size.

How can we decompose the change in magnitude of the *t*-values as the result of the inclusion of state fixed effects on the one hand, and the reduction in sample size on the other? We attempt to do this heuristically by estimating the models with the small data set and without the state fixed effects. The results of this exercise are presented at the end of Table 3. First, we observe that the coefficients associated with the marriage hazard are very similar to the coefficients in the model with fixed effects. Thus, the reduction in the magnitude of these coefficients after the introduction of fixed effects seems to be the result of the smaller sample. The results also indicate that the reduction in sample size plays an important role on the reduction in significance after the inclusion of fixed effects. In fact, approximately 63% of the observed gap between the *t*-values of models with and without state fixed effects is the result of using a subsample for the estimation of the models with fixed effects. We interpret this as strong evidence in support of the hypothesis that with a larger sample the effect of welfare on the hazard of premarital birth would be significant even in the models with fixed effects. We suspect that this may also be the case for the hazard of marriage. Also, observe that the estimated coefficients capture the marginal impact of welfare within a particular age interval, the compounded effect may be larger. We address this issue next.

*Policy Simulation Results*

From the discussion above, we are inclined to conclude that welfare had a significant effect on marriage and fertility of young Hispanic women. How important is this effect? We provide an alternative characterization of the results that will help us to appreciate the magnitude of the welfare effect on marriage and premarital birth. The characterization is based on the estimate of the average probability of marriage/premarital birth given in (5). The corresponding predicted probability by age $t$ can be obtained by replacing the expected value in (5) by the sample average, evaluated at the parameter estimates.

Table 4 reports the change in the average probability of marriage and premarital birth, respectively, as a result of a 10% percent increase in welfare generosity. The policy simulation was conducted using the estimated model 3 (series expansion with $K = 2$) with and without state fixed effects. Although the model allows us to compute the results of the policy at any age between 13 and 24, for simplicity of exposition, we present the outcomes of the policy change at four different ages, seventeen, nineteen, twenty-two and twenty-four. We observe the largest impact of welfare on the probability of premarital birth but the impact of the policy on marriage is similar in magnitude, except at age 17. We find that a 10% increase in welfare generosity would have resulted in a 10% increase in premarital births and a 7% decrease in marriages by age 24; both effects are significant. At a first glance, it may seem surprising to observe significant policy impacts given the weak significance of the individual welfare parameters. However, observe that parameters associated to welfare variables in hazard functions measure the marginal effect of welfare at a particular point in time, while the effects captured in our policy analysis measure the compounded effect of welfare after several years of exposure to welfare.

[Table 4 about here]

*Comparisons with Black and White Samples*

We discuss estimation results for the sample of white and black women. We also provide empirical evidence on parameter bias by racial group. These results are included to allow comparisons with the more detailed results for the Hispanic sample. Estimates of the parameters associated with the welfare variable and the hazards of marriage and premarital birth are reported in a web table, Table A3. For the sample of black women, the coefficient associated with the welfare variable and the premarital birth hazard has negative sign in all model specifications. This is inconsistent with the predictions of the theory. However, the coefficient is insignificant in all model specifications except for those with a small number of explanatory variables and no state fixed effects. On the other hand, the coefficient associated with the welfare variable and the marriage hazard has negative sign, as predicted by the theory. However, the coefficient is only significant for models that do not include state fixed effects.

For the sample of white women, the coefficient associated with the welfare variable and the premarital birth hazard has negative sign in all model specifications but its effect is insignificant. On the other hand, the coefficient associated with the welfare variable and the marriage hazard has negative sign, as predicted by the theory. This coefficient is significant at the usual level on all model specifications that do not include state fixed effects, and for model specifications that include state fixed effects the coefficient is significant at the ninety percent significance level, or close to it, for most model specifications. Finally, the magnitude of the coefficient does not change much across model specifications. This evidence strongly suggests that welfare has a significant effect on a white woman's decision to marry.

Overall, our results for the black and white samples are consistent with previous research. Other authors have found insignificant effects of welfare on the fertility behavior of blacks and whites (Moffitt, 1998). Our results on the effects of welfare on marriage are also consistent with existing work (Schultz, 1994).

We indicate the differences in results of the semiparametric estimation technique when compared to models that impose tighter parametric assumptions. The possibility of bias due to parametric distributional assumptions on the unobserved heterogeneity component of hazard models has already been documented. As it turns out, in our empirical application we find that distributional assumptions have the strongest effect on the baseline hazard. These results are analogous to findings in Heckman and Singer (1984). In that paper the baseline hazard in an empirical application was also affected by the choice of parametric versus non-parametric unobserved heterogeneity component. The result was a dramatic change in the shape of the baseline hazard from negative duration dependence in the parametric case to positive duration dependence in the semiparametric case.

Another web table (Table A5) presents parameter estimates for the baseline hazard under different model specifications. The baseline hazard for the Hispanic sample remains unchanged across model specifications. The same is not true for the black and white samples, with the large change observed for the black sample with a 40% average change in the fertility baseline hazard, and a 25% average change in the marriage baseline hazard, between model one (no unobserved heterogeneity) and model 4 (series expansion truncated at $K = 3$). It is also important to note that model 2, which controls for unobserved heterogeneity by means of a parametric gamma specification often used in empirical applications, displays estimates very similar to model 1.

In our application we always observe positive duration dependence across all models and samples. This is consistent with our intuition that, other things being the same, older women in our sample are more likely to marry or became pregnant than younger women. However, the flexible specification for the unobserved heterogeneity component results in a significant increase in the degree of duration dependence for whites and more so for blacks, but not for Hispanics. This suggests that a significant proportion of early marriage and premarital births in the white and black populations are due to individual specific characteristics unobserved to the econometrician, or unobserved heterogeneity, while this is not the case for the Hispanic population. In this sense, the Hispanic population constitutes a much more homogeneous group in terms of marriage and fertility behavior than the group of white, and especially black, women. Hazard models that rely heavily on parametric assumptions will fail to capture these behavioral differences across ethnic groups.

## Conclusions

The paper examined the effects of welfare generosity on a young woman's premarital fertility and marriage choices for a sample of young Hispanic women. The empirical specification considered is a competing risk model with two possible risks, the risk of premarital fertility and the risk of marriage. This specification allows us to identify the process of premarital fertility and the process of marriage. To evaluate the validity of our results we have estimated a variety of econometric specifications for different sets of explanatory variables, state fixed effects and individual specific unobserved heterogeneity.

Economic theory suggests that welfare should increase the likelihood of premarital birth. Overall, the evidence indicates that welfare has a significant effect on a Hispanic woman's decision to marry and a woman's premarital birth decisions. The direction of this effect is consistent with the predictions of the theory. Results from a comparative analysis of young women by race suggest that the Hispanic population constitutes a much more homogeneous group in terms of marriage and fertility behavior than the group of white, and specifically black, women.

A policy simulation exercise considers the average change in the probability of marriage, and premarital birth, as a result of a ten percent increase in welfare generosity. Our findings indicate that such increase would have resulted in a 10% increase in premarital births and a 7% decrease in marriages by age 24; both effects are significant.

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**TABLE 1**
**Means and Standard Deviations of Relevant Variables for Hispanic Sample**

| | **Full Sample** | **Marriage First** | **Premarital Birth** |
|---|---|---|---|
| **Family Background at Age 14** | | | |
| Mother’s Education | 8.332 (3.64) | 8.234 (3.72) | 8.318 (3.39) |
| Father’s education | 8.955 (4.09) | 8.629 (4.16) | 9.194 (3.85) |
| Family income ($ 1000’s) | 14.264 (11.7) | 13.85 (10.7) | 14.71 (12.2) |
| Number of Siblings | 4.617 (3.07) | 4.559 (2.91) | 4.449 (2.95) |
| Multiple Reading Materials at Home | 0.284 (0.45) | 0.266 (0.44) | 0.235 (0.43) |
| **Living Arrangements at Age 14** | | | |
| Both Parents | 0.678 (0.47) | 0.707 (0.46) | 0.643 (0.48) |
| Single Mother | 0.173 (0.38) | 0.159 (0.37) | 0.214 (0.41) |
| Single Father | 0.008 (0.089) | 0.003 (0.06) | 0.010 (0.10) |
| Mother and Step Father | 0.067 (0.25) | 0.062 (0.24) | 0.031 (0.17) |
| Father and Step Mother | 0.012 (0.11) | 0.007 (0.08) | 0.020 (0.14) |
| Other Family Arrangements | 0.062 (0.24) | 0.062 (0.24) | 0.082 (0.28) |
| **Religious Attendance** | | | |
| Not at all | 0.012 (0.11) | 0.013 (0.11) | 0.008 (0.09) |
| Frequently (at least once a week) | 0.411 (0.49) | 0.417 (0.49) | 0.480 (0.50) |

| | | | |
|---|---|---|---|
| **Aptitude Tests** | | | |
| Standardized AFQT Score | 0.000 ( 1.00) | -0.047 (0.92) | 0.078 (0.98) |
| **Residence at Age 14** | | | |
| Urban | 0.885 (0.32) | 0.872 (0.33) | 0.929 (0.26) |
| **Other** | | | |
| Private School | 0.065 (0.25) | 0.052 (0.22) | 0.061 (0.24) |

Note: The Income variable is measured in thousands of real dollars of 1982. Standard deviations are in parentheses.

**TABLE 2**
**Descriptive Statistics by Race and Age**

| **Average Age:** | **Hispanic** | | **Black** | | **White** | |
|---|---|---|---|---|---|---|
| First marriage | 21.31 (3.70) | | 21.31 (3.70) | | 22.12 (3.78) | |
| Premarital Birth | 19.74 (3.36) | | 19.83 (3.60) | | 20.26 (3.86) | |
| **Age/Probability of:** | **Marriage** | **Prem. Birth** | **Marriage** | **Prem. Birth** | **Marriage** | **Prem. Birth** |
| **17** | 0.082864 | 0.076309 | 0.02079 | 0.117455 | 0.047097 | 0.020621 |
| **19** | 0.218634 | 0.167871 | 0.079308 | 0.266968 | 0.183429 | 0.058505 |
| **22** | 0.424208 | 0.244601 | 0.248558 | 0.450987 | 0.429229 | 0.107437 |
| **24** | 0.513747 | 0.280737 | 0.360349 | 0.529131 | 0.57879 | 0.123144 |

**Note:** Probabilities of Marriage and Premarital Birth are obtained from Kaplan-Meier estimates. Standard deviations are in parentheses.

**TABLE 3**
**Welfare Effects across Different Model Specifications (Hispanic Sample)**

| | **Model 1** | **Model 2** | **Model 3** | **Model 4** | **Model 5** |
|---|---|---|---|---|---|
| | | | **Small Set of Explanatory Variables** | | |
| **Marriage** | -0.0005 (1.13) | -0.0007 (1.14) | -0.0006 (1.08) | -0.0006 (1.18) | -0.0006 (1.13) |
| **Pre. Fert.** | 0.0004 (0.76) | 0.0003 (0.43) | 0.0004 (0.66) | 0.0004 (0.51) | 0.0004 (0.58) |
| **LLF** | -3870.91 | -3866.34 | -3859.80 | -3863.80 | -3862.14 |
| | | | **Small Set of Explanatory Variables & State Fixed Effects** | | |
| **Marriage** | -0.0011 (0.72) | -0.0017 (0.99) | -0.0015 (0.89) | -0.0012 (0.78) | -0.0013 (0.79) |
| **Pre. Fert.** | 0.0010 (0.46) | 0.0007 (0.30) | 0.0008 (0.37) | 0.0004 (0.18) | 0.0006 (0.24) |
| **LLF** | -3115.02 | -3113.18 | -3113.27 | -3115.14 | -3114.55 |
| | | | **Large Set of Explanatory Variables** | | |
| **Marriage** | -0.0013 (2.11) | -0.0013 (2.02) | -0.0014 (1.66) | -0.0013 (1.50) | -0.0014 (1.59) |
| **Pre. Fert.** | 0.0021 (2.78) | 0.0021 (2.71) | 0.0022 (2.31) | 0.0022 (2.16) | 0.0022 (2.10) |
| **LLF** | -3838.03 | -3837.13 | -3830.96 | -3827.37 | -3823.27 |
| | | | **Large Set of Explanatory Variables & State Fixed Effects** | | |

| | | | | | |
|---|---|---|---|---|---|
| **Marriage** | -0.0008 (0.49) | -0.0009 (0.49) | -0.0009 (0.49) | -0.0009 (0.48) | -0.0010 (0.51) |
| **Pre. Fert.** | 0.0034 (1.41) | 0.0034 (1.39) | 0.0034 (1.38) | 0.0034 (1.36) | 0.0034 (1.30) |
| **LLF** | -3099.27 | -3099.13 | -3098.87 | -3098.86 | -3096.81 |
| | **Large Set of Explanatory Variables (Small Data-Set)** | | | | |
| **Marriage** | -0.0009 (1.11) | -0.0009 (1.04) | -0.0009 (1.03) | -0.0009 (0.96) | -0.0009 (0.92) |
| **Pre. Fert.** | 0.0020 (1.95) | 0.0020 (1.88) | 0.0020 (1.77) | 0.0020 (1.56) | 0.0020 (1.45) |
| **LLF** | -3109.02 | -3107.95 | -3106.20 | -3106.17 | -3105.45 |

**Note:** Model 1 does not control for unobserved heterogeneity, in model 2 the series expansion is truncated at K=1, which corresponds to the case of a bivariate gamma distribution, in models 3 to 5 we consider series expansions truncated at K=2,3 and 4 respectively. The small set of explanatory variables includes welfare generosity, the standardized AFQT test, mother's education, the number of siblings, and family income. For a list of variables included in the large set of explanatory variables, see table A2. See the results section for the rational of using the small data set. T-values are in parentheses.

**TABLE 4**
**Policy Simulation for a 10% Increase in Welfare Generosity (Hispanic Sample)**

| | **Without State Fixed effects** | | | **With State Fixed Effects** | | |
|---|---|---|---|---|---|---|
| **Age at Marriage ≤** | **Predicted Probability** | **Change in Probability** | **%Change** | **Predicted Probability** | **Change in Probability** | **%Change** |
| **17** | 0.0876 (0.048) | -0.0063 (0.0035) | -0.0763 (0.036) | 0.1063 (0.038) | -0.0055 (0.0036) | -0.0542 (0.034) |
| **19** | 0.2222 (0.065) | -0.0161 (0.0071) | -0.0758 (0.034) | 0.2332 (0.058) | -0.0138 (0.0066) | -0.0611 (0.030) |
| **22** | 0.4413 (0.101) | -0.0303 (0.0122) | -0.0725 (0.033) | 0.4276 (0.084) | -0.0282 (0.0102) | -0.0679 (0.026) |
| **24** | 0.5140 (0.108) | -0.0328 (0.0130) | -0.0680 (0.033) | 0.4969 (0.093) | -0.0338 (0.0108) | -0.0700 (0.025) |
| **Age of Prem. Birth ≤** | | | | | | |
| **17** | 0.1048 (0.065) | 0.0127 (0.0071) | 0.1283 (0.051) | 0.0741 (0.042) | 0.0158 (0.0063) | 0.2209 (0.057) |
| **19** | 0.1959 (0.075) | 0.0220 (0.0099) | 0.1167 (0.044) | 0.1410 (0.050) | 0.0279 (0.0091) | 0.2030 (0.050) |
| **22** | 0.3145 (0.088) | 0.0322 (0.0129) | 0.1062 (0.040) | 0.2281 (0.065) | 0.0409 (0.0120) | 0.1841 (0.044) |
| **24** | 0.3382 (0.089) | 0.0337 (0.0132) | 0.1035 (0.040) | 0.2430 (0.069) | 0.0422 (0.0120) | 0.1789 (0.043) |

**Note:** Policy simulation computed using Model 3 with a large set of explanatory variables. The "change in probability" represents the actual change in the value of the predicted probability as a result of the policy. The "% Change" is defined as the % change in probability with respect to the original "predicted probability." Standard deviations (in parentheses) were computed using 500 random draws from the distributions of the estimated parameters.

# APPENDIX: Additional Tables and Figures

**Figure 1: Welfare Generosity**

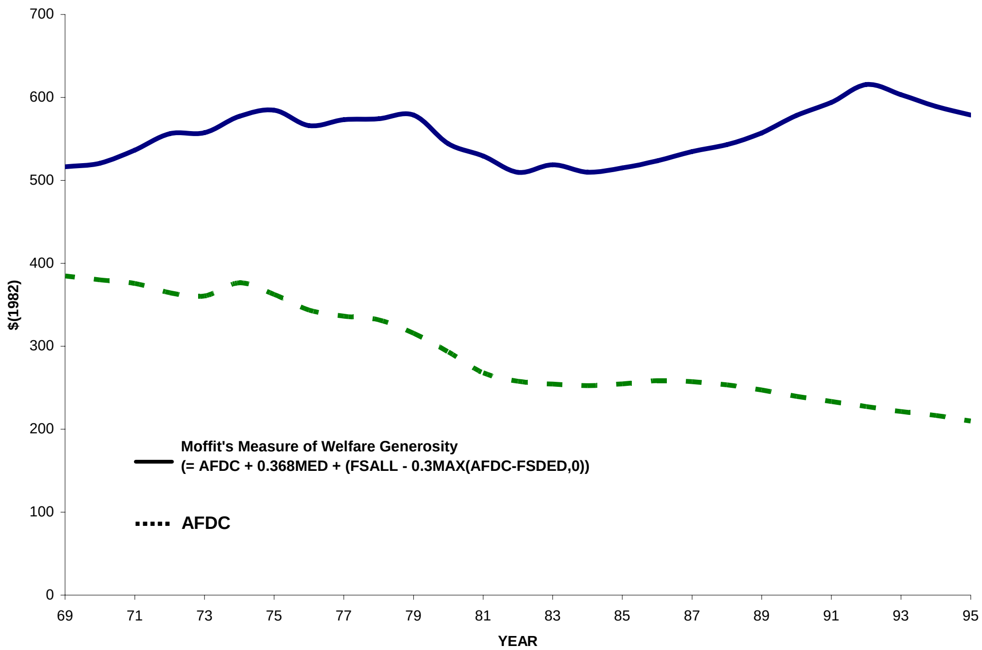

**TABLE A1**
**Estimated Parameter Values for Hispanic Sample (Models with a Small Set of Variables)**

| | Model 1 | Model 2 | Model 3 | Model 4 | Model 5 |
|---|---|---|---|---|---|
| | **Without State Fixed Effects** | | | | |
| **Marriage** | | | | | |
| Mother's Education | 0.0011 (0.07) | 0.0034 (0.14) | 0.0071 (0.30) | 0.0034 (0.19) | 0.0037 (0.18) |
| Family income ($ 1000's) | 0.0005 (0.08) | -0.0036 (0.39) | -0.0024 (0.26) | -0.0006 (0.09) | -0.0034 (0.43) |
| Number of Siblings | 0.0020 (0.11) | -0.0109 (0.38) | -0.0061 (0.22) | -0.0017 (0.08) | -0.0102 (0.43) |
| Standardized AFQT | -0.0980 (1.56) | -0.1518 (1.62) | -0.1464 (1.55) | -0.1023 (1.43) | -0.1084 (1.37) |
| Urban Residence | -0.0122 (0.08) | -0.1125 (0.51) | -0.0875 (0.40) | -0.0160 (0.09) | -0.0430 (0.23) |
| **Prem. Fertility** | | | | | |
| Mother's Education | 0.0135 (0.62) | 0.0118 (0.42) | 0.0124 (0.47) | 0.0113 (0.40) | 0.0118 (0.44) |
| Family income ($ 1000's) | -0.0072 (0.80) | -0.0103 (0.90) | -0.0092 (0.86) | -0.0109 (0.94) | -0.0120 (1.08) |
| Number of Siblings | -0.0158 (0.61) | -0.0285 (0.86) | -0.0227 (0.73) | -0.0285 (0.86) | -0.0312 (1.00) |
| Standardized AFQT | -0.0422 (0.46) | -0.0181 (0.16) | -0.0364 (0.33) | -0.0307 (0.26) | -0.0445 (0.40) |
| Urban Residence | 0.1061 (0.46) | 0.1590 (0.54) | 0.1212 (0.44) | 0.1225 (0.41) | 0.0847 (0.30) |
| | **With State Fixed Effects** | | | | |
| **Marriage** | | | | | |
| Mother's Education | -0.0080 (0.44) | -0.0041 (0.16) | -0.0054 (0.22) | -0.0069 (0.35) | -0.0071 (0.37) |
| Family income ($ 1000's) | 0.0040 (0.60) | 0.0002 (0.02) | 0.0028 (0.31) | 0.0029 (0.39) | 0.0031 (0.42) |
| Number of Siblings | 0.0009 (0.04) | -0.0113 (0.37) | -0.0064 (0.22) | -0.0036 (0.15) | -0.0035 (0.15) |
| Standardized AFQT | -0.0631 (0.87) | -0.1071 (1.05) | -0.0935 (0.97) | -0.0729 (0.90) | -0.0689 (0.88) |
| Urban Residence | -0.0382 (0.23) | -0.1121 (0.47) | -0.0910 (0.41) | -0.0514 (0.28) | -0.0490 (0.27) |
| **Prem. Fertility** | | | | | |
| Mother's Education | 0.0075 (0.27) | 0.0073 (0.23) | 0.0068 (0.22) | 0.0081 (0.25) | 0.0069 (0.19) |
| Family income ($ 1000's) | -0.0018 (0.17) | -0.0028 (0.23) | -0.0023 (0.20) | -0.0039 (0.31) | -0.0054 (0.40) |
| Number of Siblings | -0.0214 (0.66) | -0.0257 (0.71) | -0.0246 (0.71) | -0.0292 (0.80) | -0.0379 (0.96) |
| Standardized AFQT | 0.0068 (0.06) | 0.0187 (0.15) | 0.0161 (0.13) | 0.0152 (0.12) | 0.0303 (0.21) |
| Urban Residence | 0.0878 (0.34) | 0.0700 (0.24) | 0.0719 (0.26) | 0.0655 (0.22) | 0.0186 (0.06) |

**Note:** Parameter estimates associated with variables other than the welfare variable. T-values are in parentheses

**TABLE A2**
**Estimated Parameter Values for Hispanic Sample (Models with a Large Set of Variables and no State Fixed Effects)**

| | Model 1 | Model 2 | Model 3 | Model 4 | Model 5 |
|---|---|---|---|---|---|
| **Marriage** | | | | | |
| Mother's Education | -0.0103 (0.56) | -0.0102 (0.52) | -0.0113 (0.43) | -0.0102 (0.39) | -0.0109 (0.41) |
| Father's education | 0.0118 (0.71) | 0.0117 (0.66) | 0.0154 (0.66) | 0.0160 (0.69) | 0.0158 (0.66) |
| Family income ($ 1000's) | 0.0041 (0.62) | 0.0038 (0.52) | 0.0025 (0.26) | 0.0027 (0.28) | 0.0010 (0.10) |
| Number of Siblings | 0.0168 (0.82) | 0.0152 (0.69) | 0.0094 (0.32) | 0.0130 (0.44) | 0.0110 (0.37) |
| Reading Materials at Home | -0.0521 (0.42) | -0.0455 (0.34) | -0.0238 (0.13) | -0.0492 (0.28) | -0.0031 (0.02) |
| Living with Both Parents | 0.0002 (0.00) | 0.0042 (0.02) | 0.0004 (0.00) | 0.0029 (0.01) | 0.0180 (0.07) |
| Living with Single Father | -0.1190 (0.16) | -0.1274 (0.16) | -0.0525 (0.06) | -0.0757 (0.08) | -0.1364 (0.14) |
| Living with Mother/Step Father | -0.2421 (0.91) | -0.2742 (0.96) | -0.4058 (1.06) | -0.4252 (1.10) | -0.4868 (1.18) |
| Living with Father/Step Mother | -0.4117 (0.65) | -0.4506 (0.68) | -0.7575 (0.82) | -0.5833 (0.70) | -0.6304 (0.73) |
| Other Family Arrangements | 0.1283 (0.52) | 0.1498 (0.57) | 0.1632 (0.44) | 0.1499 (0.40) | 0.2005 (0.51) |
| Not religious | -1.3548 (1.95) | -1.4107 (1.97) | -1.7431 (2.06) | -1.7287 (2.03) | -1.8406 (2.00) |
| Religious | -0.0954 (0.92) | -0.0939 (0.84) | -0.0839 (0.57) | -0.0842 (0.57) | -0.0674 (0.45) |
| Standardized AFQT Score. | 0.3223 (1.52) | 0.3333 (1.46) | 0.3959 (1.32) | 0.4216 (1.41) | 0.4682 (1.47) |
| Square S. AFQT Score. | -0.1211 (2.02) | -0.1274 (2.00) | -0.1601 (1.93) | -0.1657 (2.00) | -0.1841 (2.05) |
| Urban | 0.0274 (0.18) | 0.0138 (0.08) | -0.0307 (0.14) | -0.0246 (0.11) | -0.0668 (0.30) |
| Private school | 0.0338 (0.14) | 0.0023 (0.01) | -0.0582 (0.18) | -0.0675 (0.21) | -0.1354 (0.40) |
| Real wage | -0.0077 (0.17) | -0.0124 (0.26) | -0.0452 (0.76) | -0.0514 (0.82) | -0.0460 (0.71) |
| Abortion providers | 0.3422 (4.03) | 0.3591 (3.89) | 0.4571 (3.96) | 0.4530 (3.82) | 0.4687 (3.65) |
| **Prem. Fertility** | | | | | |
| Mother's Education | -0.0021 (0.08) | -0.0031 (0.11) | -0.0043 (0.14) | -0.0043 (0.14) | -0.0049 (0.15) |
| Father's education | 0.0169 (0.75) | 0.0186 (0.79) | 0.0223 (0.81) | 0.0226 (0.81) | 0.0223 (0.78) |
| Family income ($ 1000's) | -0.0087 (0.86) | -0.0092 (0.87) | -0.0107 (0.87) | -0.0108 (0.86) | -0.0108 (0.85) |
| Number of Siblings | -0.0152 (0.53) | -0.0162 (0.55) | -0.0197 (0.57) | -0.0200 (0.57) | -0.0198 (0.55) |
| Reading Materials at Home | -0.0030 (0.02) | -0.0056 (0.03) | -0.0081 (0.04) | -0.0089 (0.04) | -0.0022 (0.01) |
| Living with Both Parents | 0.0720 (0.30) | 0.0733 (0.30) | 0.0988 (0.34) | 0.1011 (0.34) | 0.1061 (0.35) |
| Living with Single Father | 0.6086 (0.85) | 0.6165 (0.77) | 0.7271 (0.73) | 0.7342 (0.71) | 0.7556 (0.71) |
| Living with Mother/Step Father | -0.2769 (0.75) | -0.2780 (0.73) | -0.2894 (0.66) | -0.2886 (0.65) | -0.3009 (0.66) |
| Living with Father/Step Mother | 0.1534 (0.20) | 0.1298 (0.16) | 0.1285 (0.13) | 0.1307 (0.13) | 0.1374 (0.13) |
| Other Family Arrangements | 0.4139 (1.39) | 0.4250 (1.34) | 0.5141 (1.33) | 0.5207 (1.32) | 0.5355 (1.31) |
| Not religious | -0.9665 (0.85) | -1.0005 (0.86) | -1.1886 (0.89) | -1.1976 (0.88) | -1.2363 (0.87) |
| Religious | 0.1012 (0.69) | 0.1008 (0.67) | 0.1092 (0.62) | 0.1098 (0.61) | 0.1101 (0.59) |
| Standardized AFQT Score | 0.2623 (0.83) | 0.2836 (0.86) | 0.3670 (0.96) | 0.3739 (0.97) | 0.3736 (0.95) |
| Square S. AFQT Score | -0.0919 (1.05) | -0.0976 (1.07) | -0.1204 (1.14) | -0.1222 (1.14) | -0.1222 (1.12) |
| Urban | 0.0326 (0.13) | 0.0406 (0.16) | 0.0625 (0.21) | 0.0642 (0.21) | 0.0635 (0.20) |
| Private school | -0.0676 (0.23) | -0.0637 (0.20) | -0.0657 (0.18) | -0.0657 (0.18) | -0.0747 (0.19) |
| Real wage | -0.2027 (3.40) | -0.2083 (3.28) | -0.2338 (3.13) | -0.2349 (3.08) | -0.2400 (2.94) |
| Abortion providers | -0.1109 (0.97) | -0.1097 (0.93) | -0.1195 (0.82) | -0.1206 (0.72) | -0.1223 (0.73) |

**Note:** Parameter estimates associated with variables other than the welfare variable. T-values are in parentheses. The wage variable represents the real average yearly wages associated with the services, retail and wholesale trade industries at the state level. Abortion providers refer to the number of non-hospital abortion providers at the state level per 100000 women (From Guttmacher Institute).

**Table A3**
**Estimated Parameter Values for Hispanic Sample (Models with a Large Set of Variables and State Fixed Effects)**

| | Model 1 | Model 2 | Model 3 | Model 4 | Model 5 |
|---|---|---|---|---|---|
| **Marriage** | | | | | |
| Mother's Education | -0.0206 (1.01) | -0.0205 (0.99) | -0.0205 (0.97) | -0.0204 (0.96) | -0.0208 (0.90) |
| Father's education | 0.0195 (1.08) | 0.0196 (1.07) | 0.0198 (1.06) | 0.0198 (1.05) | 0.0216 (1.06) |
| Family income ($ 1000's) | 0.0036 (0.50) | 0.0036 (0.47) | 0.0035 (0.45) | 0.0035 (0.45) | 0.0035 (0.41) |
| Number of Siblings | 0.0117 (0.52) | 0.0113 (0.50) | 0.0111 (0.47) | 0.0111 (0.47) | 0.0110 (0.43) |
| Reading Materials at Home | 0.0041 (0.03) | 0.0063 (0.04) | 0.0084 (0.06) | 0.0080 (0.06) | 0.0173 (0.11) |
| Living with Both Parents | 0.0018 (0.01) | 0.0025 (0.01) | 0.0030 (0.01) | 0.0028 (0.01) | -0.0020 (0.01) |
| Living with Single Father | -0.0824 (0.12) | -0.0785 (0.11) | -0.0764 (0.11) | -0.0783 (0.11) | -0.0450 (0.06) |
| Living with Mother/Step Father | -0.1601 (0.54) | -0.1648 (0.55) | -0.1711 (0.56) | -0.1711 (0.56) | -0.1957 (0.59) |
| Living with Father/Step Mother | -0.3186 (0.50) | -0.3252 (0.51) | -0.3349 (0.51) | -0.3350 (0.51) | -0.4120 (0.56) |
| Other Family Arrangements | 0.1611 (0.60) | 0.1632 (0.59) | 0.1637 (0.58) | 0.1637 (0.55) | 0.1558 (0.48) |
| Not religious | -1.9731 (2.01) | -1.9832 (2.02) | -1.9986 (2.03) | -1.9991 (2.03) | -2.0934 (2.09) |
| Religious | -0.0201 (0.17) | -0.0190 (0.16) | -0.0179 (0.15) | -0.0177 (0.15) | -0.0190 (0.14) |
| Standardized AFQT Score. | 0.2898 (1.24) | 0.2919 (1.23) | 0.2959 (1.22) | 0.2966 (1.21) | 0.3122 (1.17) |
| Square S. AFQT Score. | -0.1085 (1.62) | -0.1098 (1.62) | -0.1116 (1.61) | -0.1118 (1.60) | -0.1194 (1.58) |
| Urban | -0.0306 (0.18) | -0.0337 (0.19) | -0.0365 (0.21) | -0.0363 (0.20) | -0.0462 (0.24) |
| Private school | -0.0969 (0.34) | -0.1044 (0.36) | -0.1098 (0.37) | -0.1095 (0.37) | -0.1058 (0.33) |
| Real wage | -0.0859 (0.93) | -0.0872 (0.93) | -0.0889 (0.92) | -0.0889 (0.91) | -0.0975 (0.93) |
| Abortion providers | 0.2535 (1.18) | 0.2567 (1.19) | 0.2620 (1.18) | 0.2621 (1.09) | 0.2892 (0.97) |
| **Prem. Fertility** | | | | | |
| Mother's Education | 0.0022 (0.07) | 0.0023 (0.07) | 0.0023 (0.07) | 0.0024 (0.07) | 0.0027 (0.07) |
| Father's education | 0.0132 (0.46) | 0.0133 (0.46) | 0.0135 (0.46) | 0.0135 (0.46) | 0.0147 (0.47) |
| Family income ($ 1000's) | -0.0033 (0.27) | -0.0034 (0.28) | -0.0035 (0.28) | -0.0036 (0.28) | -0.0048 (0.35) |
| Number of Siblings | -0.0098 (0.28) | -0.0101 (0.29) | -0.0103 (0.29) | -0.0104 (0.29) | -0.0135 (0.35) |
| Reading Materials at Home | -0.0412 (0.19) | -0.0437 (0.20) | -0.0454 (0.21) | -0.0458 (0.21) | -0.0745 (0.31) |
| Living with Both Parents | 0.1218 (0.38) | 0.1231 (0.39) | 0.1241 (0.38) | 0.1243 (0.37) | 0.1280 (0.36) |
| Living with Single Father | -0.4168 (0.21) | -0.4352 (0.22) | -0.4470 (0.21) | -0.4483 (0.21) | -0.6523 (0.25) |
| Living with Mother/Step Father | -0.1630 (0.37) | -0.1635 (0.37) | -0.1639 (0.36) | -0.1635 (0.36) | -0.1666 (0.34) |
| Living with Father/Step Mother | -0.2944 (0.24) | -0.2943 (0.24) | -0.2954 (0.24) | -0.2955 (0.24) | -0.2897 (0.22) |
| Other Family Arrangements | 0.3168 (0.82) | 0.3168 (0.80) | 0.3165 (0.79) | 0.3171 (0.75) | 0.2997 (0.65) |
| Not religious | -0.8490 (0.71) | -0.8556 (0.71) | -0.8629 (0.71) | -0.8640 (0.70) | -0.9271 (0.71) |
| Religious | 0.1344 (0.75) | 0.1349 (0.74) | 0.1352 (0.72) | 0.1352 (0.71) | 0.1447 (0.70) |
| Standardized AFQT Score. | 0.4006 (1.07) | 0.4057 (1.05) | 0.4103 (1.05) | 0.4109 (1.02) | 0.4567 (1.02) |
| Square S. AFQT Score. | -0.1097 (1.04) | -0.1110 (1.03) | -0.1122 (1.02) | -0.1125 (1.00) | -0.1235 (1.00) |
| Urban | 0.0406 (0.14) | 0.0404 (0.14) | 0.0405 (0.14) | 0.0405 (0.14) | 0.0337 (0.11) |
| Private school | 0.0117 (0.03) | 0.0136 (0.04) | 0.0152 (0.04) | 0.0150 (0.04) | 0.0433 (0.11) |
| Real wage | -0.2663 (1.84) | -0.2670 (1.82) | -0.2678 (1.80) | -0.2680 (1.75) | -0.2740 (1.70) |
| Abortion providers | -0.0776 (0.20) | -0.0755 (0.19) | -0.0728 (0.18) | -0.0724 (0.17) | -0.0600 (0.13) |

**Note:** Parameter estimates associated with variables other than the welfare variable. T-values are in parentheses. One possible explanation as to why intelligent women marry less is that younger women may have access to better options outside of marriage than the other groups of women.

**TABLE A4**
**Welfare Effects across Different Model Specifications for the Black and White Samples**

| | Model 1 | Model 2 | Model 3 | Model 4 | Model 5 |
|---|---|---|---|---|---|
| | **Black Sample** | | | | |
| | **Small Set of Explanatory Variables** | | | | |
| **Marriage** | -0.0017 (3.90) | -0.0017 (3.70) | -0.0017 (3.77) | -0.0018 (3.75) | -0.0018 (3.73) |
| **Pre. Fert.** | -0.0008 (2.06) | -0.0008 (2.06) | -0.0010 (2.26) | -0.0012 (2.36) | -0.0012 (2.36) |
| | **Small Set of Explanatory Variables & State Fixed Effects** | | | | |
| **Marriage** | -0.0016 (1.25) | -0.0016 (1.25) | -0.0017 (1.26) | -0.0017 (1.25) | -0.0017 (1.27) |
| **Pre. Fert.** | -0.0010 (0.82) | -0.0013 (0.92) | -0.0011 (0.80) | -0.0010 (0.73) | -0.0009 (0.70) |
| | **Large Set of Explanatory Variables** | | | | |
| **Marriage** | -0.0018 (3.07) | -0.0018 (3.00) | -0.0018 (2.93) | -0.0018 (2.84) | -0.0018 (2.74) |
| **Pre. Fert.** | -0.0001 (0.22) | -0.0001 (0.23) | -0.0001 (0.25) | -0.0001 (0.24) | -0.0002 (0.25) |
| | **Large Set of Explanatory Variables & State Fixed Effects** | | | | |
| **Marriage** | -0.0012 (0.78) | -0.0012 (0.76) | -0.0011 (0.71) | -0.0011 (0.66) | -0.0011 (0.67) |
| **Pre. Fert.** | 0.0000 (0.02) | 0.0000 (0.03) | 0.0003 (0.21) | 0.0006 (0.38) | 0.0007 (0.45) |
| | **White Sample** | | | | |
| | **Small Set of Explanatory Variables** | | | | |
| **Marriage** | -0.0012 (4.31) | -0.0013 (4.27) | -0.0013 (4.28) | -0.0014 (4.32) | -0.0015 (4.41) |
| **Pre. Fert.** | -0.0005 (0.88) | -0.0007 (1.10) | -0.0009 (1.39) | -0.0008 (1.24) | -0.0008 (0.00) |
| | **Small Set of Explanatory Variables & State Fixed Effects** | | | | |
| **Marriage** | -0.0012 (1.70) | -0.0013 (1.73) | -0.0013 (1.75) | -0.0014 (1.80) | -0.0014 (1.86) |
| **Pre. Fert.** | -0.0006 (0.36) | -0.0006 (0.37) | -0.0006 (0.38) | -0.0006 (0.39) | -0.0007 (0.43) |
| | **Large Set of Explanatory Variables** | | | | |
| **Marriage** | -0.0009 (2.89) | -0.0010 (2.95) | -0.0011 (2.98) | -0.0011 (3.03) | -0.0011 (3.00) |
| **Pre. Fert.** | -0.0008 (1.25) | -0.0010 (1.36) | -0.0012 (1.46) | -0.0010 (1.45) | -0.0011 (1.47) |
| | **Large Set of Explanatory Variables & State Fixed Effects** | | | | |
| **Marriage** | -0.0014 (1.54) | -0.0014 (1.55) | -0.0015 (1.59) | -0.0015 (1.60) | -0.0015 (1.62) |
| **Pre. Fert.** | 0.0000 (0.02) | 0.0000 (0.02) | 0.0000 (0.01) | 0.0000 (0.01) | 0.0000 (0.02) |

**Note:** Model 1 does not control for unobserved heterogeneity, in model 2 the series expansion is truncated at K=1, which corresponds to the case of a bivariate gamma distribution, in models 3 to 5 we consider series expansions truncated at K=2,3 and 4 respectively. The small set of explanatory variables includes welfare generosity, the standardized AFQT test, mother's education, the number of siblings, and family income. For a list of variables included in the large set of explanatory variables, see table A2. T-values are in parentheses.

**TABLE A5**
**Baseline Hazard Estimates for Several Models by Race**

| | Baseline Hazard | Model 1 | Model 2 | Model 4 | % Change M1 to M2 | % Change M1 to M4 |
|---|---|---|---|---|---|---|
| | | **Hispanic Sample** | | | | |
| **Marriage** | **13 to 17** | -4.4939 (2.71) | -4.4584 (2.66) | -4.4183 (2.62) | 0.007900 | 0.016823 |
| | **18 to 21** | -2.9988 (1.87) | -2.9546 (1.82) | -2.9009 (1.77) | 0.014739 | 0.032646 |
| | **22 to 24** | -2.7110 (1.73) | -2.6510 (1.65) | -2.5789 (1.59) | 0.022132 | 0.048727 |
| | | | | **Mean % Change:** | 0.014924 | 0.032732 |
| **Premarital Birth** | **13 to 17** | -4.7310 (2.03) | -4.7239 (2.01) | -4.7100 (1.91) | 0.001501 | 0.004439 |
| | **18 to 21** | -3.5792 (1.60) | -3.5645 (1.56) | -3.5419 (1.46) | 0.004107 | 0.010421 |
| | **22 to 24** | -4.2187 (1.91) | -4.1966 (1.83) | -4.1632 (1.70) | 0.005239 | 0.013156 |
| | | | | **Mean % Change:** | 0.003615 | 0.009339 |
| | | **Black Sample** | | | | |
| **Marriage** | **13 to 17** | -3.8249 (2.47) | -3.7736 (2.39) | -3.3480 (2.06) | 0.013412 | 0.124683 |
| | **18 to 21** | -2.1046 (1.42) | -2.0496 (1.34) | -1.5483 (0.98) | 0.026133 | 0.264326 |
| | **22 to 24** | -1.9149 (1.32) | -1.8474 (1.22) | -1.2283 (0.78) | 0.035250 | 0.358557 |
| | | | | **Mean % Change:** | 0.024932 | 0.249188 |
| **Premarital Birth** | **13 to 17** | -4.1216 (2.65) | -4.0509 (2.55) | -3.0058 (1.71) | 0.017154 | 0.270720 |
| | **18 to 21** | -3.2724 (2.19) | -3.1755 (2.07) | -1.8880 (1.11) | 0.029611 | 0.423053 |
| | **22 to 24** | -3.4966 (2.37) | -3.3732 (2.20) | -1.6856 (0.99) | 0.035291 | 0.517932 |
| | | | | **Mean % Change:** | 0.027352 | 0.403902 |
| | | **White Sample** | | | | |
| **Marriage** | **13 to 17** | -5.1365 (5.92) | -5.0529 (5.66) | -4.5017 (4.79) | 0.016276 | 0.123586 |
| | **18 to 21** | -3.3026 (3.97) | -3.2103 (3.71) | -2.6182 (2.87) | 0.027948 | 0.207231 |
| | **22 to 24** | -2.9512 (3.65) | -2.8369 (3.29) | -2.0908 (2.30) | 0.038730 | 0.291542 |
| | | | | **Mean % Change:** | 0.027651 | 0.207453 |
| **Premarital Birth** | **13 to 17** | -4.9611 (2.12) | -4.9247 (2.05) | -4.3135 (1.69) | 0.007337 | 0.130536 |
| | **18 to 21** | -4.0182 (1.78) | -3.9739 (1.71) | -3.2802 (1.31) | 0.011025 | 0.183664 |
| | **22 to 24** | -4.3151 (1.95) | -4.2619 (1.85) | -3.4240 (1.37) | 0.012329 | 0.206507 |
| | | | | **Mean % Change:** | 0.010230 | 0.173569 |

**Note:** Model 1 does not control for unobserved heterogeneity, in model 2 the series expansion is truncated at K=1, which corresponds to the case of a bivariate gamma distribution, model 4 considers a series expansions truncated at K=3. All models include a large set of explanatory variables and state fixed effects. For a list of variables included in the large set of explanatory variables, see table A2. T-values are in parentheses.