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BY

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Abstract

The minimum legal drinking age (MLDA) requirements can affect teen fertility rates through changes in alcohol-induced risky sexual behavior. The direction of the effect can vary depending on changes in alcohol consumption context and intensity. Using micro-level data, I find that a decrease in the MLDA increases the probability of unwanted first pregnancy among 15-20 year-old blacks and poor whites. The effect on non-poor whites is not statistically significant. I find some evidence that the individual eligibility status at the time of first pregnancy rather than the state MLDA might affect fertility among non-poor whites.

JEL classification code: J13; J18

Keywords: Minimum Legal Drinking Age (MLDA), Pregnancy, Fertility, Sexual Behavior, Alcohol Consumption, Discrete Hazard

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I. Introduction

Numerous studies associate teenage alcohol consumption with increased motor vehicle accident mortality, sexually transmitted diseases, date rape, and other risk-taking behaviors with long-term consequences.[1] I analyze the effect of alcohol consumption restrictions measured in terms of the state minimum legal drinking age (MLDA) on teen pregnancy, birth and abortion rates. Understanding the causes of teen childbearing is crucial for designing effective public policy. As noted by Fletcher and Wolfe (2009), among other adverse consequences, teen childbearing lowers human capital accumulation for teen mothers and affects their earnings.

The causal effect of teen alcohol use on teen pregnancy is thought to be mediated through risky sexual behavior (Dee 2001 and Sen 2003). Several studies report that sexually active teens under the influence of alcohol are less likely to use contraception (Markowitz et al. 2005) and are consequently more likely to experience an unintended pregnancy. If easy availability of alcohol leads to a higher likelihood of unintended pregnancy, then this increase in the number of pregnancies could result in a relatively larger increase in the number of abortions than live births, as unintended pregnancies are more likely to be terminated than planned pregnancies (Finer and Henshaw 2006). Strict restrictions on legal alcohol availability, on the other hand, should be associated with a decrease in the number of pregnancies, births, and/or abortions. This hypothesis so far has received mixed empirical support. Dee (2001) reports that an increase in the MLDA to 21 years – which represents a decrease in the availability of alcohol – reduces state level birth rates among black 15-19 year-old teens, but does not affect their white peers.

Some studies suggest that risky sexual behavior depends on the intensity of alcohol intake per unit of time and the context in which the alcohol is consumed (O'Hare 2005). For example, moderate alcohol consumption in a bar has different implications than binge drinking at a home party. Legal restrictions that limit adolescent freedom to consume alcohol might create a rebellious response expressed in increased efforts to acquire alcohol and binge drinking instead of moderate drinking.[2] If more restrictions on alcohol consumption increase the likelihood of more intense drinking in the private setting where intimacy is more likely to occur, then there could be a relatively large number of unintended pregnancies and abortions in the presence of a

[1] For more details, see review studies by Wagenaar and Toomey (2002) and Newbury-Birch (2009).

[2] Allen et al. (1994) presents a review of the literature on this topic.

high MLDA and a decrease when restrictions are relaxed. These are the opposite effects that many policy-makers might expect.

I test the relationship between the drinking eligibility restrictions and the fertility of young women using exogenous variations in the MLDA across states in the 1970s and 1980s.[3] There are several contributions my paper introduces to the existing alcohol/teenage fertility literature. Unlike a few related studies that examine the relationship between alcohol consumption restrictions and fertility of 15-19 year-old teens using state level aggregate data, I use micro-level data from the 1979 cohort of the National Longitudinal Survey of Youth (NLSY). The NLSY micro-level panel allows for the reconciliation of state-level alcohol policy variables with individual fertility variables such as occurrence of first pregnancy, allowing me to explore heterogeneous effects by individual characteristics. Such precision is not possible with annually aggregated fertility data. Secondly, in contrast to the literature focused only on one pregnancy outcome at a time (i.e., births or abortions), I model the effect on the probability of first pregnancy and the probability of birth, abortion or miscarriage of that pregnancy. The focus on first incidence is novel and reasonable, given that the target group of MLDA is teens and most of them who get pregnant experience this as a first time event. Finally, I analyze the effect on different race-age groups which in some instances is not feasible with aggregate data due to peculiarities of abortion statistics that are not systematically reported by state, year, race, and age.

Following the existing literature, I limit my analysis to the MLDA restrictions for beer as it is the most popular alcoholic beverage among youth (Coate and Grossman 1988).[4] Prior to 1988, the MLDA restrictions were regulated at the state-level creating considerable variation across states, with the lowest age requirement set at 18 and the highest at 21 years of age. Between 1970 and 1975, the number of states with a MLDA below 21 years increased from 18 to 39 states. However, under the threat of losing Federal highway funding, by July 1988 the MLDA was raised and set uniformly across the United States to 21. The changes in the MLDA for beer by

[3] Reduction in the MLDA across country in the early 1970s was alongside a decrease in the voting age; increases in the MLDA in the 1980s were compelled by the federal law requirements.
[4] If binge drinking is driven by non-beer alcoholic beverages then the MLDA for beer might not be an appropriate measure. During 1970-1988, only 15 states had stricter age restrictions for liquor; all other states had the same age for both. With an exception of conditional probability models that rely on small samples, the analysis of the MLDA for distilled spirits yields qualitatively similar results.

state from 1970 to 1990 are summarized in Appendix A. Additionally, in 1973, after the Supreme Court ruling in *Roe v. Wade,* abortions became legal on the national level, significantly reducing the cost of unintended pregnancy for women nationwide.[5]

The empirical analysis relies on a discrete-time hazard model using the micro-level monthly fertility data. Two measures of eligibility restrictions are used in separate analyses: an indicator of whether a woman resides in a state with the MLDA below 21 years and an indicator of whether a woman can legally drink, given her age and the MLDA in her state of residence. The former variable is consistent with the literature; the latter variable provides a more precise measure of person-specific eligibility that accounts for the "grandfather" clause not captured by the MLDA dummy. Using predicted probabilities, I evaluate the effect of the change (decrease or increase) in the MLDA and the effect of becoming legally eligible to drink on the probability of becoming pregnant for the first time separately for 15-20, 15-17, 18-20, and 21-23 year-old white, black, and Hispanic women. Given the information on how the first pregnancy ended and the MLDA restrictions at the time of pregnancy conception, I evaluate the effect of changes in the eligibility restrictions on the probability of first pregnancy ending in a live birth, an abortion or a miscarriage. Identifying possible effects of the changes in the MLDA is of high priority in light of recent debate regarding the effects of the MLDA.[6]

The results indicate that a decrease in the MLDA to 18, 19, or 20 years significantly increases the probability of becoming pregnant for the first time among black and poor white 15-20 year-old women (respectively by 0.5 and 0.9 percentage points). This is a substantial effect given that the base monthly probability for this age group is 1.1% and 0.7% respectively. Among Hispanic women, the probability of first pregnancy tends to be higher in states where the MLDA is set to 21 years. Consequently, a decrease in the MLDA reduces the incidence of first pregnancy. As for non-poor white women, changes in the MLDA do not affect pregnancy rates in a substantive economic or statistical way. In addition, a decrease in the MLDA tends to have a substantial positive effect on the probability of termination of first pregnancy among pregnant

[5] The abortion laws are discussed in Gold (2003). The "pill", introduced in the 1960s, might have affected fertility rates. However, despite being the most effective method of contraception used by the sexually active teens in the 1970s, it also had a high misuse rate (see Jones and Darroch Forrest 1989).
[6] In recent years, the Amethyst Initiative, a public movement calling for a re-examination of the MLDA of 21, was supported by the leadership on many campuses.

15-20 year old black, poor white, and Hispanic women. This might be consistent with an increase in the number of unwanted pregnancies among blacks and poor-whites in states with the MLDA below 21. The discrete hazard model of first pregnancy passes a falsification test, as I do not find an association between the MLDA and the probability of first pregnancy among women age 21-23.

The stratification into 15-17 and 18-20 age groups reveals that the effect on the probability of first pregnancy is generally robust only for 18-20 year olds, which is the age group directly affected by the MLDA changes. In comparison, the estimates for 15-17 year old women are much smaller and statistically insignificant. Given that the effect is driven by the behavior of women who are directly affected by changes in age restrictions (i.e., 18-20 year olds), the individual eligibility, rather than the state's age limit, might be a better measure of the implied effect of alcohol consumption restrictions. I find that, compared to their non-eligible peers, non-poor white women who can drink legally have lower probability of becoming pregnant for the first time and also lower probabilities of abortion or miscarriage of the first pregnancy, indicating that there are fewer unwanted pregnancies among women who can drink legally. This finding is consistent with a proposition that reaching the legal age for alcohol consumption expands the set of activities for 18-20 year-olds and opens an alternative less risky venue for alcohol consumption (i.e. drinking at the bar) to binge drinking at someone's house.[7] The results for poor-whites and blacks, that rely on much smaller samples, are mixed and the conditional probabilities must particularly be interpreted with caution.

Overall, the results support both hypothesized mechanisms. The direction of the effect that the MLDA has on the probability of first pregnancy among women age 15-20 varies across population sub-groups, and it might indicate that the mechanism that links alcohol consumption patterns and risky sexual behavior differs across groups of women as well. I find some evidence that the MLDA below 21 increases incidence of unwanted pregnancies among some women. For other women, the eligibility status at the time of first pregnancy that can affect the context in which the alcohol is consumed, rather than the state MLDA at the time of first pregnancy, might be a stronger determinant of fertility. This could mean that the MLDA alone might not be an

[7] Wechsler et al. (2000) reports some evidence in favor of this hypothesis: the underage students tend to drink less often, but more drinks per occasion; they are also more likely to drink in private settings.

adequate measure to capture alcohol consumption behavior and one should account for the intensity, location, and legality of alcohol consumption as well.

II. Literature Review: Alcohol and Teen Fertility

Teen demand for alcohol is relatively responsive to changes in alcohol consumption restrictions. Most studies conclude that a decrease in restrictions on alcohol consumption, as well as a decrease in taxation that affects prices of alcoholic beverage, lead to an increase in teen alcohol consumption (Grossman et al. 1994, Coate and Grossman 1988, Dee and Evans 2003).[8]

Despite a positive association between alcohol use and risky sexual behavior among teenagers reported in numerous studies, the causal nature of the relationship remains unknown.[9] Results drawn primarily from aggregate data can only assume that occasions of alcohol use and risky sex coincide; event specific studies, in contrast, provide some evidence in favor of an association for young heterosexuals (Donovan and McEwan 1995). Some studies report that alcohol use may lower contraception use among sexually active teens, and, therefore, increase the probability of an unplanned pregnancies (e.g., Grossman and Markowitz 2005; Markowitz et al. 2005; Hingson et al. 1990, and Rees et al. 2001).

A few recent studies based on aggregate data have emphasized the relationship between alcohol consumption restrictions and teen fertility. Using a "difference-in-difference-in-difference" model, Dee (2001) finds that the nationwide increases in the MLDA to 21 reduced the birth rate among black 15-19 year-old teens by roughly 5.5 percent; the effect on white teens is mostly statistically insignificant and "implausibly" signed. Dee suggests that underlying differences in patterns of sexual behavior and alcohol consumption could explain variation in racial childbearing patterns and race-specific responses to changes in drinking age. Sen (2003) investigates the effects of beer taxes and other alcohol-related policies, including the MLDA, on teen pregnancy outcomes (abortion and birth). The MLDA does not appear to have a robust, statistically significant impact on either outcome. One should be cautious with the interpretation

[8] For exceptions see Dee (1999) and Kaestner (2000). The literature review on the topic is summarized in Wagenaar et al. (2009) and Wagenaar and Toomey (2002).
[9] See literature review on subject in Leigh and Stall (1993) and Rashad and Kaestner (2004).

of these results, as Sen's study relies on four years of data (1985, 1988, 1992, and 1996), and only the period from 1985 to 1988 involves variation in the MLDA.[10]

Both Dee (2001) and Sen (2003) use state-level panel data on birth and abortion rates. The main disadvantage of this approach in studying the effect of MLDA on teen fertility is that it does not permit a thorough analysis of the effects of external factors at the time of pregnancy on individual decision making. Results presented by Dee (2001) only partially support the hypothesis that a low MLDA leads to higher childbearing rates among teenagers and indicate that the response differs across races. Additionally, aggregate data on abortions and pregnancies might not be informative for this kind of analysis for three reasons. First, abortion data are not available by state, year, age group, and race. Therefore, unlike Dee (2001), one neither can construct abortion and pregnancy rates by race nor estimate separate models for each race. Second, a lack of uniform reporting requirements across states introduces a substantial number of missing values in the abortion data collected by the Center for Disease Control and Prevention (CDC). This affects the quality of pregnancy rates, calculation of which relies on abortion data. Third, CDC's abortion data reflect abortions by the state where abortion took place and not by the state of residence. The latter is more preferred if one wants to study the effect of changes in the state's law on the rates in the same state.[11]

III. Empirical Analysis

Since aggregate data on childbearing make it difficult to study the policy effects that differ across race-age groups, a more appropriate policy analysis should rely on disaggregated data which permit a separation of the potential effects by race and by age of the individual. I capture the relationship between alcohol consumption eligibility restrictions and teen fertility using the micro-level monthly fertility data that allows me to incorporate individual characteristics in the analysis as well as control for the state policy at the time of first pregnancy.

A. Data

I use the 1979 cohort of the NLSY that consists of a nationally representative random sample of young men and women who were 14 to 22 years old in 1979 and oversamples of

[10] A sensitivity test reported in the same study indicates that other discussed results are not robust to the choice of covariates (e.g., models with state versus region fixed effects).
[11] The abortion statistics collected by the Alan Guttmacher Institute (AGI) is by state of residence. However, their data are not available by age group.

young blacks, Hispanics, poor whites, and members of the military. The use of NLSY is advantageous as it contains detailed retrospective fertility and mobility histories, which permit tracking the timing of the first pregnancy and its outcome (birth, abortion, miscarriage/stillbirth) in the 1970s-1980s for 6,283 women and identifying the state of residency and the state MLDA for each woman at the time of that pregnancy.[12]

The identification of pregnancy incidence relies on the reported number of pregnancies and their outcomes. Abortions are underreported in the NLSY data, especially in the earlier survey years (Jones and Darroch Forrest 1992). Udry et al. (1996), report that blacks and Hispanics are significantly less approving of abortion in a variety of circumstances than whites, and these differences translate into different propensities to report. I address this potential issue by estimating models separately for non-black, non-Hispanic women (whites), non-Hispanic blacks (blacks), and Hispanics.

The timing of events becomes of high priority when one attempts to study the effect of the change in the state policy on individual decisions. To make my analysis as precise as possible, I combine retrospective information and data obtained from the 1979-1988 annual surveys and convert the NLSY data set into a panel where a unit of observation is a person-month. Each woman enters my dataset in the month when she turns 15 years old. For every month after the entry, I know whether she became pregnant or not, her state of residence, and the MLDA in that state. Once she turns 21 years old she exits the data set, as past this age drinking age restrictions are not binding.[13] Further restrictions include exclusion of women with incomplete fertility history, women serving in military, and women who became pregnant before their 15th birthday. The final sample includes 399,528 monthly observations during the period 1972-1985 on 5,549 women. About 60% of my sample is white women, one-quarter is black, and the rest are Hispanics.

The person-month data set can be viewed as transition data where women can move from one state (being not pregnant) to another (becoming pregnant). Table 1 presents a narrative history of the pregnancy occurrence and its outcome for women age 15-20 and for comparison

[12] The National Survey of Family Growth (NSFG) contains a larger sample and provides more details on contraception, abortion, and sexual behavior but no data on the retrospective mobility history. Therefore, the timing of past pregnancies cannot be matched to the state of residence or the state policy.
[13] For a falsification test, an extended dataset includes 21-23 year olds.

for women age 21-23 described in terms of hazard and survival functions. The examination of survival probabilities reveals that 95% of 5,549 women at risk of first pregnancy did not become pregnant at age 15, 80% – did not experience a first pregnancy by their 18th birthday, and slightly more than half of the sample did not have it by their 21st birthday. About 75% of all first pregnancies that occurred to 15-20 year-old women ended in a live birth and 16% were terminated.[14]

The race specific incidence shows that 54% of the black sample became pregnant for the first time between ages 15 to 20 years; the pregnancy occurrence among poor whites and Hispanics was 47% compared to 32% for women in the non-poor white sample. The inspection of composition of pregnancies aggregated by race, age, and the state MLDA, reported in supplementary Appendix C, reveals that the timing of first pregnancies differs across races as well. For example, the majority of all first time pregnant 15, 16, and 17 year-olds were black women (the respective shares of 42%, 35%, and 32%); among all 18-20 year-olds nearly 35% were non-poor whites and only 27% were blacks. The share of pregnancies among poor whites and Hispanics are rather stable across age groups. An additional stratification by the MLDA shows that a disproportionally higher share of first pregnancies among non-poor whites, poor whites, and especially blacks occurs in states where the MLDA is set below 21. (Similarly, a higher share of first pregnancies occurs among 18-20 year-old black, poor white, and non-poor white women who are legally eligible to drink). In terms of pregnancy resolution, the share of pregnancies that are terminated appears to be higher in states where the MLDA is below 21 years, but only for poor white and Hispanic samples.

B. Model Specification: Discrete-Time Hazard

The fertility history data identify the month and year when first pregnancy occurs. This suggests grouping observations into discrete (monthly) time intervals that cover period from January 1972 through December 1985. I use the discrete-time hazard model which accommodates the nature of the observed data and the research question. Vast literature exists on

[14] The birth ratio among women pregnant for the first time is 5% points higher than the corresponding statistics from the NSFG surveys; the abortion ratio is 1.5% point higher (supplementary Appendix B). These numbers cannot be compared to the national statistics as the latter reflects a share of all pregnancies that have been terminated. Abortion ratios among blacks and Hispanics age 15-20 are slightly lower compared to the NSFG surveys; for white - about the same.

discrete-time models of event history data (e.g., Allison 1982; Singer and Willett 2003); here, I briefly discuss the main concept of these models.

Becoming pregnant for a first time at time *t* is a non-repeatable event which is intrinsically conditional on not experiencing the event at any time period prior to *t*. Let *T* denote the discrete random variable whose values T_i indicate the time period *t* when the i^{th} woman experiences her first pregnancy. The conditional probability h_{ist} that a randomly selected woman *i* in state *s* will experience her first pregnancy in time *t*, given a set of observable characteristics *X*, is defined as

$$h_{ist} = \Pr[T_i = t \mid T_i \geq t, X_{ist}], \quad (1)$$

For each woman the dependent variable that indicates whether she is pregnant for the first time can be represented as a string of zeros (indicating not pregnant) followed by a one (indicating pregnancy). If a woman did not have the first pregnancy prior to age 21 then the dependent variable is represented only by a string of zeros implying that the event has yet to be experienced. The binary nature of the dependent variable and the specification of the log-likelihood function allow one to model the hazard probability as a probit function.[15]

The presence of several age cohorts in my data set determines two notions of time: calendar time and age-time. For each woman I observe the age when she had her first pregnancy which also corresponds to a certain calendar time *t*. I normalize age by expressing it in terms of months since birth minus 180 so it corresponds to months since age 15. To accommodate both notions of time, in addition to the index for calendar time *t*, I introduce the age counter τ that represents the time that a woman spends at risk of first pregnancy measured in months (subject to normalization discussed above). To allow for flexibility, the function *g(τ)* is parameterized as a cubic polynomial and it captures the effect of age (and the duration dependency) on the hazard probability after accounting for other covariates. The effect of calendar time is captured by a full set of calendar time fixed effects.

Empirically the discrete-time (monthly) hazard probability of first pregnancy is given by the following specification

$$\Pr[1^{st}\_pregnancy_{ist} = 1 \mid \text{not pregnant before t}, X_{ist}] =$$
$$\Phi(\beta_0 + \delta\, \mathbf{eligibility\_restriction_{ist}} + \beta' X_{ist} + \theta_1 age_{it} + \theta_2 age_{it}^2 + \theta_3 age_{it}^3 + \gamma_s + \eta_t), \quad (2)$$

[15] Allison (1982) shows that the log-likelihood function for the discrete-time hazard model is the log-likelihood function for the regression analysis of dichotomous dependent variables. It is a common practice to use the logit or probit link functions in estimation of discrete-time hazard models.

where 1^{st}_pregnancy is a binary indicator of the pregnancy status, Φ – the standard cumulative normal distribution function, and i indexes individuals, s – state of residence, t – calendar time that corresponds to a combination of month and year and ranges from April 1972 to September 1985. For estimation purpose, η_t corresponds to a set of calendar year and calendar month dummies. The latter is included to allow for a seasonality effect. Standard errors are clustered by state.

I use two measures of eligibility restrictions (two separate models) to capture the effect of the MLDA on teen fertility: a dummy indicating whether the MLDA in a state of residence is set below age 21 and an indicator of whether a teen can legally drink in the state of residence. The latter variable might be helpful in identifying channels through which the MLDA restrictions affect teen behavior. A significant number of states increased the MLDA while allowing a "grandfather clause," a provision that exempts teens who were previously eligible to drink from new eligibility requirements. If a female is "grandfathered" by the law, the "legally eligible" dummy reflects this minutia. Table 2 provides the descriptive statistics for variables included in the model. Over 70% of women age 15-20 in the data set were residing in states with the MLDA below 21 years, but only one third of them can drink legally.

The vector X included in (2) contains controls for individual and family characteristics: race, religion in which female was raised, the Armed Forces Qualifications Test (AFQT) scores, whether both parent were present in the household at age 14, mother's education, and an indicator for the marital status. The use of contraception before the first pregnancy, which is included in the model, is an important determinant of fertility, but it contains a large number of non-response values (one third of women included in the data set) resulting in a significant reduction of sample sizes. Other considered controls, such as the presence of an older sibling, father's education, an indicator for the enforced parental involvement laws did not improve models' fit and, therefore, are omitted from the model (2).

To test whether the effect of eligibility restrictions varies across age groups and races, I estimate separate equations for white, black, and Hispanic 15-20 year-old women. Additionally, whites are disaggregated into poor white and non-poor white subsamples. The segmentation of the data permits all estimated coefficients to vary across race groups. The preference is given to separate equations rather than a single equation model with numerous interaction terms since the interpretation of the interaction terms in a nonlinear model is not straightforward, as the sign and the magnitude of the effect vary with the values of covariates (Ai and Norton 2003). Finally, alcohol consumption patterns among 15-17 year olds, who are not legally eligible to drink under any MLDA regime, might differ from patterns among 18-20 year olds, who are directly affected by the MLDA variations. To take this into account, I estimate models by race separately for 15-17 and 18-20 year olds. I test the credibility of the model by repeating the estimation for women age 21-23. If the model is correct, then eligibility restrictions should not affect the dependent variable for this age group, as it is not subject to the MLDA restrictions.

Overall, the eligibility restrictions could have a positive or a negative impact on the probability of first pregnancy. The former will be observed if there is a complimentary relationship between alcohol consumption and risky sex that increases instances of unplanned pregnancies; the latter – if easy availability of alcohol affects the place and intensity of alcohol consumption toward binge drinking in the private setting where risky sex is more likely to occur. Also, as suggested by Fertig and Watson (2009), the MLDA restrictions might affect the composition of the pool of women who become pregnant. Since I know the outcome of the first pregnancy (birth, abortion or miscarriage) and the alcohol consumption eligibility restrictions in a state at the time of conception of that pregnancy, I can explore the compositional aspect by estimating the probability of abortion, the probability of a live birth, and the probability of a miscarriage among women age 15-20 who are pregnant for the first time.[16] These models are estimated as the conditional on pregnancy probability models using a specification similar to (2). Calendar month dummies in these specifications are omitted; calendar year and state fixed effects are included.

Empirical results are discussed in the next section. Since probit estimates per se do not provide meaningful information regarding the magnitude of estimated effects (Greene 1998), the general discussion is omitted here (estimates are shown in supplementary Appendix D). Instead, the discussion is focused on the "average marginal effects" calculated using a discrete first-difference approach. The effects of personal characteristics on studied probabilities have generally expected signs, but as they are of secondary interest the discussion is omitted.

C. Results

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[16] Further disaggregation into 15-17 and 18-20 age groups substantially reduces sample sizes and produces noisy point estimates that lack credibility.

The discrete change method refers to a change in the individual predicted probability of outcome due to a change in the dichotomous explanatory variable (Long 1997). For example, the effect of a decrease in the MLDA for a woman *i* in a given race-age group (Δ_i), equals the difference between the predicted probability of first pregnancy while in a state with the MLDA below 21 years and the predicted probability of first pregnancy while in a state with the MLDA set to 21:

$$\Delta_i = \text{predicted Pr}[\text{1st\_pregnancy}_{ist} = 1 | X_{ist,} \& \text{ MLDA=18, 19 or 20}] - \text{predicted Pr}[\text{1st\_pregnancy}_{ist} = 1 | X_{ist,} \& \text{ MLDA= 21}] \quad (3)$$

Then the effect is averaged across observations in a given race-age group: $\overline{\Delta} = \sum_{i=1}^{N_R} \Delta_i / N_R$, where N_R is the number of women in a race-age group *R*. Note that the change Δ_i is calculated for each woman given her characteristics; the MLDA, on the other hand, is exogenously assigned a certain value. The latter implies that I predict probabilities for two policy scenarios: initially the MLDA is set to below 21 for everyone in the sample then it is set to 21 years for everyone in the sample. To identify the effect of an increase in the MLDA one should flip the sign of Δ_i. In a similar manner, I calculate the average change in the predicted probability of first pregnancy due to becoming legally eligible to drink. The robust standard errors for the average effects are obtained using a delta-method.

The change in the predicted probability of first pregnancy due to a decrease in the MLDA by race-age group is reported in Table 3. Blacks and poor whites tend to have lower probabilities of becoming pregnant for the first time when the MLDA is set to 21. A decrease in MLDA to 18 or 19 or 20 years significantly *increases* average monthly probability of first pregnancy by 0.9 percentage points among poor whites. The increase in the probability for Blacks is smaller (0.5 percentage points) and only marginally statistically significant. In contrast, Hispanics tend to have higher probability of becoming pregnant for the first time when the state MLDA is set to 21 and, hence, lowering of the MLDA is associated with a 1.2 percentage point *reduction* in the probability. The state MLDA has no significant effect on pregnancy probabilities among non-poor whites.

The disaggregation of 15-20 year old sample into 15-17 and 18-20 age groups provides more insights on who is affected by the MLDA. In case of poor whites, a statistically significant positive change in the probability of first pregnancy due to a decrease in the MLDA is observed for 18-20 year-old women. Yet, the effect is not statistically significant for their peers age 15-17. Similarly, for Hispanics a statistically significant decrease in the probability is observed only for 18-20 year old women (although the magnitude is rather large). For 15-17 and 18-20 year old blacks the estimated effect is positive but noisy. The effect for their white peers is negligible and also statistically insignificant. As expected, changes in the MLDA do not affect first childbearing of 21-23 year-old women.

The effects of MLDA on pregnancy resolution are reported in Table 4. A low MLDA at the time of first pregnancy does not affect that pregnancy's outcome for 15-20 year-old non-poor white sample, but significantly increases the (conditional) probability of abortion among poor white women (a change of 23.4 percentage points) and blacks (a change of 9.7 percentage points). In the case of first time pregnant Hispanic women, the MLDA below 21 is associated with the lower probability of live birth and higher probabilities of abortion and miscarriage. The decrease in the MLDA yields a 30.4 percentage point decrease in the probability of giving birth and increases the probability of abortion and miscarriage by 20.9 and 15.1 percentage points respectively. Qualitatively, the results are robust to the exclusion of a control for contraceptive use that allows one to increase the size of estimation sample.

The results described above do hint that the MLDA below 21 years tend to be associated with an increase in the number of unwanted pregnancies (as revealed by the increased incidence of abortion) among some groups of women. However, being in a state with the MLDA below age 21 is a relatively noisy measure of implied alcohol consumption that does not account for the fact that some teens, who currently live in a state with the MLDA set to 21, are "grandfathered" in the dated restrictions. Also not all 18-19 year-old teens in a state with the MLDA below 21 years can legally drink. Carpenter and Dobkin (2009), using the 1997-2005 National Health Interview Surveys, find a large and immediate increase in drinking among young adults after becoming legally eligible to drink. Therefore, the results from a specification that includes an indicator for whether a woman can legally drink, which are reported in Table 5, might be more informative.

Being legally eligible to drink at the time of pregnancy significantly increases the probability of becoming pregnant for the first time among 18-20 year-old poor white women by 0.8 percentage point. Yet, the eligibility status does not appear to have a statistically significant effect on how that pregnancy ends (i.e., live birth, abortion or miscarriage). The estimated average marginal effects have substantial standard errors, but a low precision is expected given the sample size. Among non-poor white women age 18-20 becoming legally eligible to drink alcohol yields a 0.5 percentage point decrease in the probability of first pregnancy. The same event is associated with an increase in the probability of giving birth and a decrease in the probability of abortion or miscarriage among pregnant women. As for black women, those legally eligible to drink tend to have lower probability of becoming pregnant for the first time. However, among those 18-20 year-old blacks who did become pregnant, women who are legally eligible to drink have higher probability of abortion. Results from this model are mostly in line with the ones reported for 15-20 year old blacks and poor whites. In contrast, results for non-poor whites differ substantially. Results are not reported for Hispanics as the sample size is not sufficient to produce meaningful estimates.

Results presented here cannot be easily compared to those reported in Dee (2001), since I study the hazard of first pregnancy and its outcome using individual level data and Dee studies state level birth rates that are not limited to the first time pregnant women. Nonetheless, similar to Dee, I do not find an association between changes in the MLDA and the fertility of some white women. In contrast, I find that an increase in the MLDA reduces the probability of first pregnancy and the probability of terminating the first pregnancy among 15-20 year old poor whites. As for black women age 15-20, I find that an increase in the MLDA decreases the probability of first pregnancy, but no evidence that it also reduces the birth rates among first time pregnant black women.

IV. Conclusion

Changes in the MLDA across states in the 1970s and 1980s can alter the alcohol-induced risky sexual behavior. Conventional wisdom suggests that, everything else the same, one will observe higher pregnancy, birth and/or abortion rates in states with a relatively low MLDA and the opposite when the MLDA is high. The evidence presented in the literature, primarily drawn from the aggregate state level data, indicates that an increase in the MLDA to 21 years affects only black teen birth rates, but not the birth rate of their white peers. Yet, changes in the MLDA may affect the choice of the location where alcohol consumption takes place and the intensity of alcohol intake. For example, the relatively high MLDA could induce binge-drinking among the underage which usually takes place at the private gatherings where risky sexual behavior (including unprotected sex) is more likely to occur. If so, one might observe higher pregnancy, birth, and/or abortion rates after an increase in the MLDA. Results reported in Naimi et al. (2003) suggest that implications of binge drinking may also vary across races.

I use a discrete-time hazard model that relies on individual level data from the NLSY to test whether easing alcohol availability, measured in terms of a decrease in the MLDA below 21 years and a change in an individual legal alcohol consumption status, leads to a change in teen pregnancy rate or pregnancy outcome. The contribution of this paper to the existing literature is its focus on the incidence of first pregnancy among young women. The majority of teens who get pregnant experience this as a first time event. Therefore, identifying factors that affect teen childbearing is a topic of high importance. In addition, individual level data allows me to reconcile state level alcohol policy variables at the time of the pregnancy with individual fertility decisions. This cannot be done accurately with the annual state aggregates. Finally, with the exception of birth rates, the stratification by age-race groups is also not feasible with state level fertility rates.

The results from the discrete-time hazard model indicate that a decrease in the MLDA from 21 years to 18, 19, or 20 years significantly increases the probability of first pregnancy among 15-20 year-old poor white and black women. In contrast, Hispanics tend to have lower probability of becoming pregnant for the first time when the MLDA is below 21 years. The effect on non-poor whites is trivial in both economic and statistical senses. A decrease in the MLDA tends to have a substantial positive effect on the probability of termination of the first pregnancy among 15-20 year old women (specifically blacks, poor whites, and Hispanics). This might be consistent with a conventional wisdom that states with the MLDA below 21 have a higher number of unwanted pregnancies at least among some population sub-groups.

The stratification of sample into 15-17 and 18-20 age groups shows that effect of MLDA on the incidence of first pregnancy is driven by the 18-20 age group. However, the indicator for the MLDA can be an inferior measure of the implied alcohol consumption as not all 18-20 year olds can legally drink when the MLDA is set below 21. I find that, compared to their peers who cannot legally drink, non-poor white women who can drink legally tend to have lower probability of becoming pregnant for the first time as well as lower probability of abortion or miscarriage, indicating that there are fewer unwanted pregnancies among women who can drink legally. This result is consistent with a suggestion that the pattern of alcohol consumption behavior might be triggered by changes in the eligibility status rather than the MLDA for some population sub-groups. The results for poor-whites and blacks, that rely on much smaller samples, are mixed. Given that the results differ across considered outcomes, measures of eligibility restrictions, and sub-groups of population, more research is needed to identify the mechanisms through which alcohol consumption restrictions affect teen childbearing rates.

References

Ai C, Norton E.C. "Interaction Terms in Logit and Probit Models." *Economic Letters* 2003; 80; 123-129

Allen D.N, Sprenkel D.G, Vitale P.A. "Reactance Theory and Alcohol Consumption Laws: Further Confirmation Among Collegiate Alcohol Consumers." *Journal of Studies on Alcohol* 1994; 55; 34-40

Allison P.D. "Discrete-time Methods for the Analysis of Event Histories." *Sociological Methodology* 1982; 13; 61-98

Carpenter C, Dobkin C. "The Effect of Alcohol Access on Consumption and Mortality: Regression Discontinuity Evidence from the Minimum Drinking Age." *American Economic Journal: Applied Economics* 2009; 1(1); 164-182

Coate D, Grossman M. "Effects of Alcoholic Beverage Prices and Legal Drinking Ages on Youth Alcohol Use." *Journal of Law and Economics* 1988; 31(1); 145-171

Dee T.S. "State Alcohol policies, Teen Drinking and Traffic Fatalities." *Journal of Public Economics* 1999; 72; 289-315

Dee T.S. "The Effects of Minimum Legal Drinking Ages on Teen Childbearing." *Journal of Human Resources* 2001; 36(4); 823-838

Dee T.S, Evans W.N. "Teen Drinking and Educational Attainment: Evidence from Two-sample Instrumental Variables Estimates." *Journal of Labor Economics* 2003; 21(1); 178-209

Donovan C, McEwan R. "A Review of the Literature Examining the Relationship Between Alcohol Use and HIV-related Sexual Risk-taking in Young People." *Addiction* 1995; 90; 319-328

Fertig A.R, Watson T. "Minimum Drinking Age Laws and Infant Health Outcomes." *Journal of Health Economics* 2009; 28; 737-747

Finer L.B, Henshaw S.K. "Disparities in Rates of Unintended Pregnancy in the United States, 1994 and 2002." *Perspectives on Sexual and Reproductive Health*, 38(2), 2006, 90-96

Fletcher J. M, Wolfe B. L. Education and Labor Market Consequences of Teenage Childbearing: Evidence Using the Timing of Pregnancy Outcomes and Community Fixed Effects. *Journal of Human Resources* 2009; 44(2); 303-325

Gold R.B. "Lessons from before Roe: Will Past be Prologue?" Guttmacher Report on Public Policy, 2003 March; 8-11

Greene W. “Gender Economics Courses in Liberal Arts Colleges: Further Results.” *Journal of Economic Education* 1998; 29(4); 291-300

Greenberger M.D, Connor K. “Parental Notice and Consent for Abortion: Out of Step with Family Law Principles and Policies.” *Family Planning Perspectives* 1991; 23(1); 31-35

Grossman M, Chaloupka F.J, Saffer H, Laixuthai A. “Effects of Alcohol Price Policy on Youth: A Summary of Economic Research.” *Journal of Research on Adolescence* 1994; 4(2); 347-364

Grossman M, Markowitz S. “I Did What Last Night? Adolescent Risky Sexual Behaviors and Substance Use.” *Eastern Economic Journal* 2005; 3(3); 383-405

Haas-Wilson D. “The Economic Impact of State Abortion Restrictions on Minors’ Demand for Abortions.” *Journal of Human Resources* 1996; 31(1); 140-158

O’Hare T. “Risky Sex and Drinking Contexts in Freshman First Offenders.” *Addictive Behaviors* 2005; 30(9); 585-588

Hingson R. W, Strunin L, Berlin B. M, Heeren T. “Beliefs about AIDS, Use of Alcohol and Drugs, and Unprotected Sex Among Massachusetts Adolescents.” *American Journal of Public Health* 1990; 80; 295-299

Jones E.F, Darroch Forrest J. “Contraceptive Failure in the United States: Revised Estimates from the 1982 National Survey of Family Growth.” *Family Planning Perspectives* 1989; 21; 103-109

Jones E.F, Darroch Forrest J. “Underreporting of Abortion in Surveys of U.S. Women: 1976 to 1988.” *Demography* 1992; 29(1); 113-126

Kaestner R. “A Note on the Effect of Minimum Drinking Age Laws on Youth Alcohol Consumption.” *Contemporary Economic Policy* 2000; 18(3); 315–325

Leigh B, Stall R. “Substance Use and Risky Sexual Behavior for Exposure to HIV, Issues in Methodology, Interpretation and Prevention.” *American Psychologist*, 1993, 1035-1044

Long J.S. *Regression Models for Categorical and Limited Dependent Variables. Advanced Quantitative Techniques in the Social Sciences.* Sage Publication, 1997

O’Malley P.M, Wagenaar A.C. “Minimum Drinking Age Laws: Effects on American Youth 1976-1987.” Monitoring the Future Occasional Paper Series 28, 1990

Markowitz S, Kaestner R, Grossman M. “An Investigation of the Effects of Alcohol Consumption and Alcohol Policies on Youth Risky Sexual Behaviors.” *American Economic Review* 2005; 95(2); 263-266

Merz J.F, Jackson C.A, Klerman J.A. “A Review of Abortion Policy: Legality, Medicaid Funding, and Parental Involvement, 1967-1994.” *Women's Rights Law Reporter*, 1995, 1-61

Naimi T.S, Lipscomb L.E, Brewer R.D, Gilbert B.C. “Binge Drinking in the Preconception Period and the Risk of Unintended Pregnancy: Implications for Women and Their Children.” *Pediatrics* 2003; 111; 1136-1141

New M.J. “Analyzing the Effects of State Legislation on the Incidence of Abortion During the 1990s.” The Heritage Center for Data Analysis Report 4(1), 2004

Newbury-Birch D. Impact of alcohol consumption on young people: a systematic review of published reviews. Research Report DCSF-RR067, 2009

Rashad I, Kaestner R. “Teenage Sex, Drugs and Alcohol Use: Problems Identifying the Cause of Risky Behavior.” *Journal of Health Economics* 2004; 23; 493-503

Rees D.I, Argys L.M, Averett S.L. “New Evidence on the Relationship Between Substance Use and Adolescent Sexual Behavior”. *Journal of Health Economics* 2001; 20; 835–845

Sen B. “Can Beer Taxes Affect Teen Pregnancy? Evidence Based on Teen Abortion Rates and Birth Rates.” *Southern Economic Journal* 2003; 70(2); 328-343

Singer J.D, Willett J.B. *Applied Longitudinal Data Analysis: Modeling Change and Event Occurrence.* Oxford University Press, 2003

Udry J.R, Gaughan M, Schwingl P.J, van den Berg B.J. “A Medical Record Linkage Analysis of Abortion Underreporting.” *Family Planning Perspectives* 1996; 28(5); 228-231

Wagenaar A.C. “Legal Minimum Drinking Age Changes in the United States, 1970-1981.” *Alcohol Health and Research World* 1981; 6(2); 21-26

Wagenaar A.C, Salois M.J, Komro K.A. “Effects of Beverage Alcohol Price and Tax Levels on Drinking: A Meta-analysis of 1003 Estimates from 112 studies.” *Addiction* 2009; 104; 179-190

Wagenaar A.C, Toomey T.L. “Effects of Minimum Drinking Age Laws: Review and Analyses of the Literature from 1960 to 2000.” *Journal of Studies on Alcohol Supplement* 2002; 63(2); 206-226

Wechsler H, Kuo M, Lee H, Dowdall G.W. “Environmental Correlates of Underage Alcohol Use and Related Problems of College Students.” *American Journal of Preventive Medicine* 2000; 19(1); 24-29

Table 1 – Timing of the first pregnancy and its outcome aggregated by years of age for 15-20 and 21-23 year-old women

Age in years	Teens at risk	Pregnant for the 1st time	Pregnancy ended in a live birth	Pregnancy ended in an abortion	Hazard probability of the 1st pregnancy	Survival probability of the 1st pregnancy
15	5,549	257	201	44	0.05 (0.00)	0.95 (0.00)
16	5,292	406	324	57	0.08 (0.00)	0.88 (0.00)
17	4,886	438	326	75	0.09 (0.00)	0.80 (0.01)
18	4,448	473	346	72	0.11 (0.00)	0.72 (0.01)
19	3,975	459	347	71	0.12 (0.01)	0.63 (0.01)
20	3,516	329	230	52	0.09 (0.00)	0.57 (0.01)
Women for whom the MLDA is not binding anymore						
21	3,187	314	234	40	0.10 (0.01)	0.52 (0.01)
22	2,873	276	223	35	0.10 (0.01)	0.47 (0.01)
23	2,597	251	185	31	0.10 (0.01)	0.42 (0.01)

Note: Standard errors are in the parentheses.

Table 2 – Descriptive statistics for 15-20 year-old women, 1972-1985

Variable	Mean	SD	Observations: person-month	Observations: person
Time variant variables				
Legally eligible to drink (adjusted for the grandfather clause)	0.32	0.47	395,036	5,544
MLDA is 18, 19 or 20 years	0.72	0.45	395,036	5,544
Currently married	0.11	0.31	399,528	5,549
Enforced parental notification/consent law for abortion for minors	0.02	0.13	395,757	5,549
Used contraception before 1^{st} pregnancy	0.44	0.50	275,904	3,832
Time invariant variables				
Non-poor White	0.40	0.49	399,528	5,549
Poor white	0.19	0.39	399,528	5,549
Black	0.25	0.43	399,528	5,549
Hispanic	0.17	0.37	399,528	5,549
Raised in Baptist family	0.29	0.45	398,376	5,533
Raised in other religion	0.33	0.47	398,376	5,533
Raised as Atheist	0.03	0.18	398,376	5,533
Raised in Catholic family	0.35	0.48	398,376	5,533
AFQT score below the mean	0.55	0.50	381,600	5,300
Two-parent household at age 14	0.68	0.47	398,880	5,540
Mother's education (years)	10.77	3.19	378,288	5,254
Father's education (years)	10.85	4.00	343,080	4,765
Presence of older siblings	0.78	0.41	376,344	5,227

Note: Data sources for the state MLDA: Wagenaar (1981), O'Malley and Wagenaar (1990), and the National Highway Traffic Safety Administration website. If a person resided outside of the USA then for the corresponding time periods missing values are assigned for alcohol consumption eligibility restrictions and state dummies. Data sources for the parental notification laws: Merz et al. (1995), Haas-Wilson (1996), Greenberger and Connor (1991), New (2004), and NARAL website.

Table 3 – The average predicted probability of first pregnancy and the average effect of a **decrease** in the MLDA, by race-age groups

	Non-poor White	**Poor White**	**Black**	**Hispanic**	**Non-poor White**	**Poor White**	**Black**	**Hispanic**
		15-20 age group				**21-23 age group**		
Prob. of *1st pregnancy* if MLDA=**21**	0.010	0.007	0.011	0.020	0.016	0.018	0.034	0.023
Prob. of *1st pregnancy* if MLDA<**21**	0.008	0.017	0.017	0.008	0.019	0.022	0.024	0.021
Discrete change due to a decrease in the MLDA ($\bar{\Delta}$)	**-0.002**	**0.009****	**0.005**$^{+}$	**-0.012***	**0.003**	**0.003**	**-0.010**	**-0.002**
	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Number of observations	70,007	32,982	41,288	30,406	17,157	6,016	6,317	5,906
		15-17 age group				**18-20 age group**		
Prob. of *1st pregnancy* if MLDA=**21**	0.008	0.010	0.010	0.010	0.015	0.006	0.013	0.061
Prob. of *1st pregnancy* if MLDA<**21**	0.005	0.011	0.014	0.007	0.012	0.029	0.023	0.010
Discrete change due to a decrease in the MLDA ($\bar{\Delta}$)	**-0.003**	**0.001**	**0.004**	**-0.003**	**-0.002**	**0.023****	**0.010**	**-0.051**$^{+}$
	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.01)	(0.01)	(0.03)
Number of observations	40,852	19,912	26,573	18,187	28,994	12,269	14,450	11,466

Note: ** significant at 1%; * at 5%; $^{+}$ at 10%. Robust standard errors, obtained by the delta-method, are in parentheses. Clustering of the standard errors on state produces qualitatively similar results. Standard errors for the monthly average predicted probabilities are available on request. All regressions include controls for religion in which a female was raised, the AFQT scores, mother's education, whether both parents were present in the household at age 14, indicators for being currently married and the use of contraception before the first pregnancy, a cubic polynomial for age as well as state, calendar year, and calendar month fixed effects.

Table 4 –The average conditional probability of birth, abortion, and miscarriage for 15-20 year-old women and the average effect of a **decrease** in the MLDA, by race

	Non-poor White			**Poor White**		
	Live birth	Abortion	Miscarriage	Live birth	Abortion	Miscarriage
Cond. probability of outcome if MLDA=**21**	0.580	0.317	0.123	0.536	0.038	0.188
Cond. probability of outcome if MLDA<**21**	0.663	0.233	0.135	0.813	0.271	0.102
Discrete change due to a	**0.083**	**-0.084**	**0.011**	**0.277**	**0.234****	**-0.086**
decrease in the MLDA ($\bar{\Delta}$)	(0.09)	(0.09)	(0.04)	(0.22)	(0.06)	(0.15)
Sample size		605			430	

	Black			**Hispanic**		
	Live birth	Abortion	Miscarriage	Live birth	Abortion	Miscarriage
Cond. probability of outcome if MLDA=**21**	0.848	0.067	0.158	0.863	0.083	0.067
Cond. probability of outcome if MLDA<**21**	0.793	0.164	0.098	0.559	0.293	0.218
Discrete change due to a	**-0.055**	**0.097***	**-0.060**	**-0.304****	**0.209***	**0.151***
decrease in the MLDA ($\bar{\Delta}$)	(0.05)	(0.05)	(0.10)	(0.11)	(0.09)	(0.06)
Sample size		631			359	

Note: ** significant at 1%; * at 5%; + at 10%. Robust standard errors, obtained by the delta-method, are in parentheses. Clustering of the standard errors on state produces qualitatively similar results. Standard errors for the conditional monthly average predicted probabilities are available on request. All regressions include controls for religion in which a female was raised, the AFQT scores, mother's education, whether both parents were present in the household at age 14, indicators for being currently married and the use of contraception before the first pregnancy, a cubic polynomial for age as well as state and calendar year fixed effects. Further disaggregation into 15-17 and 18-20 age groups substantially reduces sample sizes and produces noisy point estimates.

Table 5 – The average effect of **becoming legally eligible** to drink on the average probability of first pregnancy and the conditional probabilities of birth, abortion and miscarriage, 18-20 year-old women

	Non-poor White			
	1st pregnancy	Live birth	Abortion	Miscarriage
Av. probability of outcome if **cannot legally drink**	0.016	0.430	0.406	0.306
Av. probability of outcome if **can legally drink**	0.011	0.741	0.162	0.119
Discrete change due to becoming legally eligible to drink ($\bar{\Delta}$)	**-0.005*** (0.00)	**0.311**** (0.07)	**-0.244**** (0.09)	**-0.187*** (0.08)
Sample size	28,994	376		

	Poor White			
	1st pregnancy	Live birth	Abortion	Miscarriage
Av. probability of outcome if **cannot legally drink**	0.013	0.779	0.118	0.126
Av. probability of outcome if **can legally drink**	0.020	0.725	0.342	0.250
Discrete change due to becoming legally eligible to drink ($\bar{\Delta}$)	**0.008**$^{+}$ (0.00)	**-0.054** (0.16)	**0.224** (0.15)	**0.124** (0.08)
Sample size	12,269	210		

	Black			
	1st pregnancy	Live birth	Abortion	Miscarriage
Av. probability of outcome if **cannot legally drink**	0.024	0.811	0.062	0.159
Av. probability of outcome if **can legally drink**	0.019	0.766	0.210	0.156
Discrete change due to becoming legally eligible to drink ($\bar{\Delta}$)	**-0.005**$^{+}$ (0.00)	**-0.045** (0.08)	**0.148**** (0.06)	**-0.003** (0.08)
Sample size	14,450	288		

Note: ** significant at 1%; * at 5%; $^{+}$ at 10%. Robust standard errors, obtained by the delta-method, are in parentheses. Clustering of the standard errors on state produces qualitatively similar results. Standard errors for the average predicted probabilities are available on request. All regressions include controls as described in notes under Table 3 and Table 4 respectively for discrete hazard model and the conditional models for pregnancy outcomes. Results for Hispanics rely on a small sample and are not reported as the point estimates are prohibitively large and not reliable.

Appendix A: State MLDA for beer, 1970-1988

State	on1/1/1970	Changes in MLDA in the period 1/1/1970 –1/1/1988
Alabama	21	7/1/1975 - 19; 10/1/1985 - 21[gc]
Alaska	21	9/1/1970 - 19; 11/1/1984 - 21
Arizona	21	8/1/1972 - 19; 1/1/1985 - 21[gc]
Colorado	18	7/1/1987 - 21
Connecticut	21	10/1/1972 -18; 7/1/1982 -19; 10/1/1983-20; 9/1/1985-21[gc]
Delaware	21	7/1/1972 - 20; 1/1/1984 - 21[gc]
DC	18	9/1/1986 - 21[gc]
Florida	21	7/1/1973 - 18; 10/1/1980 - 19; 7/1/1985 - 21[gc]
Georgia	21	7/1/1972 -18; 9/1/1980 -19; 9/1/1985 -20; 9/1/1986 - 21
Hawaii	20	3/1/1972 - 18; 10/1/1986 - 21
Iowa	21	4/1/1972-19; 7/1/1973-18; 7/1/1976 -19[gc]; 9/1/1986 - 21[gc]
Idaho	20	7/1/1972 - 19; 4/1/1987 - 21
Illinois	21	9/1/1973 - 19; 1/1/1980 - 21
Kansas	18	7/1/1985 - 21[gc]
Louisiana	18	3/1/1987 - 21
Maine	20	6/1/1972 - 18; 10/1/1977 - 20; 7/1/1985 - 21[gc]
Maryland	21	7/1/1974 -18; 7/1/1982 - 21[gc]
Massachusetts	21	3/1/1973 - 18; 4/1/1979 - 20; 6/1/1985 - 21[gc]
Michigan	21	1/1/1972 - 18; 12/1/1978 - 21
Minnesota	21	6/1/1973 - 18; 9/1/1979 - 19[gc]; 9/1/1986 - 21[gc]
Mississippi	18	10/1/1986 - 21
Montana	21	7/1/1971 -19; 7/1/1973 -18; 1/1/1979 -19; 4/1/1987- 21
Nebraska	20	6/1/1972 - 19; 7/1/1980 - 20[gc]; 1/1/1985 - 21[gc]
New	21	6/1/1973 - 18; 5/1/1979 - 20; 6/1/1985 - 21[gc]
New Jersey	21	1/1/1973 - 18; 1/1/1980 - 19[gc]; 1/1/1983 - 21[gc]
New York	18	12/1/1982 - 19; 12/1/1985 - 21
North	18	10/1/1983 - 19; 9/1/1986 - 21
Ohio	18	8/1/1982 - 19; 7/1/1987 - 21[gc]
Oklahoma	18	9/1/1983 - 21
Rhode Island	21	3/1/1972 -18; 7/1/1980 -19; 7/1/1981 -20; 7/1/1984 - 21
South	18	1/1/1984 - 19; 1/1/1985 - 20; 9/1/1986 - 21
South Dakota	19	7/1/1972 - 18; 7/1/1984 - 19; 4/1/1988 - 21
Tennessee	21	5/1/1971 - 18; 6/1/1979 - 19; 8/1/1984 - 21[gc]
Texas	21	8/1/1973 - 18; 9/1/1981 - 19; 9/1/1986 - 21
Vermont	21	7/1/1971 - 18; 7/1/1986 - 21[gc]
Virginia	21	7/1/1974 - 18; 7/1/1981 - 19; 7/1/1985- 21[gc]
West Virginia	18	7/1/1983 - 19; 7/1/1986 - 21
Wisconsin	18	7/1/1984 - 19; 9/1/1986 - 21[gc]
Wyoming	21	5/1/1973 - 19; 7/1/1988 - 21

Note: MLDA in Arkansas, California, Indiana, Kentucky, Missouri, Nevada, New Mexico, North Dakota, Oregon, Pennsylvania, Utah, and Washington was 21 years during 1/1/1970-1/1/1988.
Source: Wagenaar (1981), O'Malley and Wagenaar (1990), and the National highway traffic safety administration (NHTSA) online reports.
[gc] Indicates that the law change includes a "grandfather clause."

Appendix B: Comparison of abortion data across NLSY and NSFG surveys

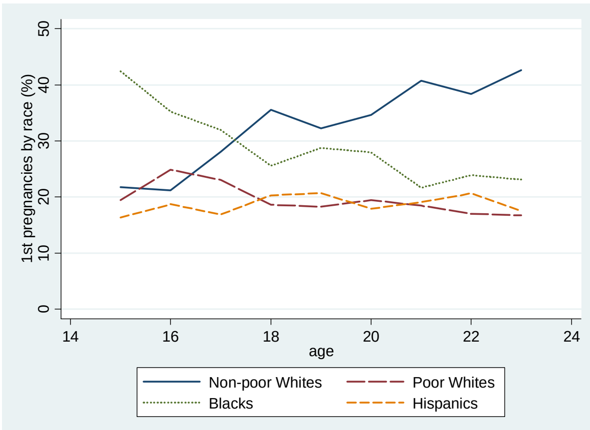
Panel A	**NLSY**					**NSFG** [a]				
Age at 1st pregnancy	Total # of 1st pregn.	1st pregn. ended in: Birth	1st pregn. ended in: Abortion	Birth ratio	Abortion ratio	Total # of 1st pregn.	1st pregn. ended in: Birth	1st pregn. ended in: Abortion	Birth ratio	Abortion ratio
15	257	201	44	78.2	17.1	726	483	164	66.5	22.6
16	406	324	57	79.8	14.0	1041	731	182	70.2	17.5
17	438	326	75	74.4	17.1	1318	934	223	70.9	16.9
18	473	346	72	73.2	15.2	1473	1023	268	69.5	18.2
19	459	347	71	75.6	15.5	1334	957	214	71.7	16.0
20	329	230	52	69.9	15.8	1165	846	149	72.6	12.8
21-23	841	642	106	76.3	12.6	2752	2021	336	73.4	12.2
Average 15-20				**75.2**	**15.8**				**70.2**	**17.3**
Average 15-19				**76.2**	**15.8**				**69.8**	**18.2**
National statistics 15-19 (AGI) [b]				**51.6**	**34.6**				**51.6**	**34.6**

Panel B	**NLSY**			**NSFG** [a]		
	Whites [c]	*Blacks*	*Hispanics*	*Whites*	*Blacks*	*Hispanics*
Total # 1st pregnancies among 15-20	1183	737	442	3970	2860	227
of which ended in birth (%)	69.1	82.6	78.5	65.8	76.8	73.1
of which ended in abortion (%)	20.5	9.5	13.1	20.6	12.2	14.5
Total # 1st pregnancies among 21-23	488	192	161	1947	687	118
of which ended in birth (%)	73.2	77.6	84.5	73.1	74.2	73.7
of which ended in abortion (%)	15.0	9.9	8.7	12.8	10.5	12.7

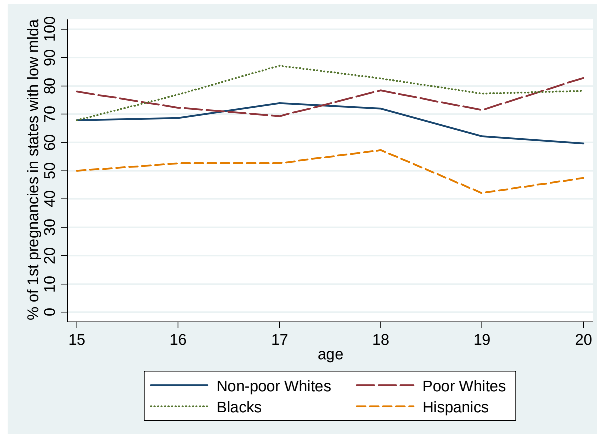
Note: [a] Pooled 1976, 1982, 1988, 1995 years. For comparability, the following restrictions are applied to the NSFG surveys: 1) only first pregnancies and their outcomes are considered (reflect 33-37% of all pregnancy records); 2) age at conception is 15-23; 3) conception occurred between 1972 and 1988. Due to restrictions 2) and 3), between 40-50% of all first pregnancy records from the NSFG 1982-1995 surveys are used; for the 1976 survey this number is 14%. [b] Averages for 1972-1985 time period. This number is not directly comparable to the corresponding NLSY or NSFG statistics as it reflects the share of *all* terminated pregnancies among women age 15-19 not the share of first pregnancies that are terminated. [c] Includes a poor white sample.

Appendix C: Composition of first pregnancies by age, race, and the state MLDA

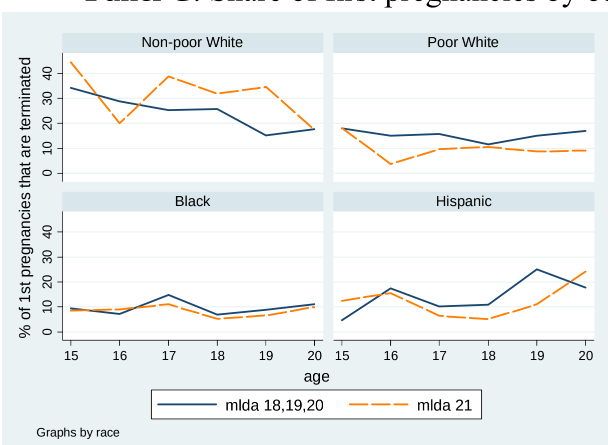
Panel A: Distribution of first pregnancies by age and race

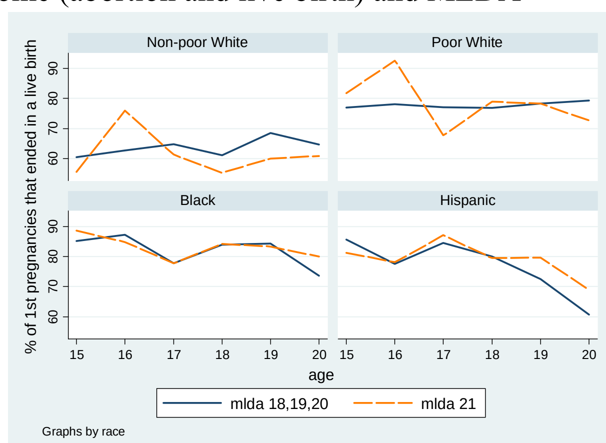


Panel B: Share of first pregnancies occurred in states with MLDA 18, 19 or 20 years



Panel C: Share of first pregnancies by outcome (abortion and live birth) and MLDA





Appendix D: Probit estimates

Table D-1: Probit coefficient estimates for 15-20 and 21-23 age groups

	15-20 year-old				*21-23 year-old*			
	Non-poor White	**Poor White**	**Black**	**Hispanic**	**Non-poor White**	**Poor White**	**Black**	**Hispanic**
MLDA is 18, 19 or 20	-0.08	0.33^{**}	0.16	-0.38^{*}	0.09	0.075	-0.16	-0.03
	(0.07)	(0.10)	(0.10)	(0.15)	(0.14)	(0.13)	(0.15)	(0.19)
Baptist family	0.05	0.10	0.05	0.33^{**}	0.08	-0.23	-0.20	-0.05
	(0.05)	(0.08)	(0.05)	(0.06)	(0.11)	(0.16)	(0.13)	(0.12)
Other religion	-0.02	0.02	0.10^{+}	0.07	0.09	-0.16	-0.23^{*}	-0.09
	(0.04)	(0.08)	(0.05)	(0.14)	(0.07)	(0.11)	(0.10)	(0.24)
Atheist family	0.15^{*}	0.24^{**}	0.14	-0.18	0.25	-0.46^{*}	0.16	0.50^{*}
	(0.06)	(0.08)	(0.10)	(0.20)	(0.19)	(0.23)	(0.29)	(0.22)
AFQTscore below the mean	0.11^{**}	0.13^{*}	0.10^{*}	0.10^{+}	0.08	0.37^{**}	0.33^{**}	0.22^{*}
	(0.04)	(0.06)	(0.04)	(0.06)	(0.06)	(0.08)	(0.11)	(0.11)
Mother's education (years)	-0.03^{**}	-0.01	-0.03^{**}	0.00	-0.02^{+}	0.01	-0.01	-0.01
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)	(0.01)
Two-parent house at 14	-0.08^{*}	-0.02	-0.19^{**}	-0.14^{**}	-0.07	-0.05	-0.01	-0.11
	(0.03)	(0.06)	(0.04)	(0.03)	(0.07)	(0.11)	(0.09)	(0.14)
Currently married	0.66^{**}	0.69^{**}	0.58^{**}	0.92^{**}	0.57^{**}	0.61^{**}	0.40^{**}	0.65^{**}
	(0.05)	(0.07)	(0.11)	(0.07)	(0.04)	(0.10)	(0.11)	(0.07)
Used contraception before 1st pregn.	-0.16^{**}	-0.17^{**}	-0.14^{**}	-0.22^{**}	0.13^{*}	0.29^{**}	0.09	-0.12^{+}
	(0.04)	(0.04)	(0.04)	(0.05)	(0.06)	(0.10)	(0.06)	(0.06)
Constant	-2.58^{**}	-2.94^{**}	-2.38^{**}	-2.61^{**}	-52.06^{**}	-35.98	28.99	-59.65^{+}
	(0.26)	(0.30)	(0.21)	(0.31)	(16.55)	(34.37)	(25.87)	(33.68)
Observations	70,007	32,982	41,288	30,406	17,157	6,052	6,317	5,906

Note: **significant at 1%; * at 5%; $^{+}$ at 10%. Standard errors clustered by state. Person-month observations. All models include state, calendar year, and calendar month fixed effects as well as a cubic polynomial for age. Excluded category for religion is women raised in Catholic families. Dependent variable: pregnancy status equals 1 if pregnant, 0 otherwise.

Appendix D (continued)

Table D-2: Probit coefficient estimates for 15-17 and 18-20 age groups

	15-17 year-old				18-20 year-old			
	Non-poor White	Poor White	Black	Hispanic	Non-poor White	Poor White	Black	Hispanic
MLDA is 18, 19 or 20	-0.18	0.05	0.13	-0.16	-0.07	0.68**	0.24	-0.90**
	(0.13)	(0.13)	(0.11)	(0.25)	(0.09)	(0.20)	(0.19)	(0.32)
Baptist family	0.12	0.02	0.17+	0.34*	-0.02	0.18+	-0.02	0.29**
	(0.10)	(0.11)	(0.09)	(0.13)	(0.07)	(0.10)	(0.10)	(0.11)
Other religion	0.02	0.02	0.27**	-0.01	-0.04	-0.00	-0.03	0.19
	(0.06)	(0.10)	(0.10)	(0.17)	(0.04)	(0.11)	(0.11)	(0.20)
Atheist family	0.26**	0.13	0.26+	0.13	0.02	0.40**	0.05	
	(0.09)	(0.16)	(0.14)	(0.29)	(0.10)	(0.13)	(0.18)	
AFQTscore below the mean	0.15*	0.10	0.04	0.08	0.09+	0.16*	0.17+	0.07
	(0.06)	(0.08)	(0.06)	(0.11)	(0.05)	(0.07)	(0.09)	(0.05)
Mother's education (years)	-0.02	-0.01	-0.03**	0.00	-0.04**	-0.01	-0.02	0.00
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Two-parent house at 14	-0.02	-0.02	-0.21**	-0.11*	-0.12**	-0.04	-0.17**	-0.17**
	(0.06)	(0.08)	(0.04)	(0.05)	(0.04)	(0.07)	(0.06)	(0.06)
Currently married	0.82**	0.82**	1.07**	1.28**	0.63**	0.67**	0.56**	0.84**
	(0.11)	(0.08)	(0.21)	(0.08)	(0.05)	(0.09)	(0.10)	(0.12)
Contrac. before 1st pregnancy	-0.38**	-0.29**	-0.26**	-0.33**	-0.01	-0.05	0.03	-0.14*
	(0.07)	(0.06)	(0.06)	(0.08)	(0.04)	(0.07)	(0.06)	(0.06)
Constant	-2.51**	-2.69**	-2.63**	-2.95**	-7.12*	-9.91+	-7.01+	-5.30
	(0.33)	(0.32)	(0.24)	(0.58)	(3.04)	(5.41)	(4.02)	(3.66)
Observations	40,852	19,912	26,573	18,187	28,994	12,269	14,450	11,466

Note: ** significant at 1%; * at 5%; + at 10%. Standard errors clustered by state. Person-month observations. All models include state, calendar year, and calendar month fixed effects as well as a cubic polynomial for age. Excluded category for religion is women raised in Catholic families. Dependent variable: pregnancy status equals 1 if pregnant, 0 otherwise.

Appendix D (continued)

Table D-3: Probit coefficient estimates for conditional on pregnancy models, women age 15-20

	Non-poor White			Poor White			Black			Hispanic		
	B	**A**	**M**	**B**	**A**	**M**	**B**	**A**	**M**	**B**	**A**	**M**
MLDA is 18,	0.28	-0.37	0.06	1.00	2.50**	-0.45	-0.26	0.69	-0.33	-1.36*	1.44*	0.87**
19 or 20	(0.29)	(0.38)	(0.24)	(0.81)	(0.84)	(0.71)	(0.26)	(0.42)	(0.49)	(0.58)	(0.67)	(0.33)
Baptist family	0.08	0.15	-0.28	0.37	0.33	-0.69^{+}	-0.31	0.23	0.40	-0.16	0.55	0.00
	(0.17)	(0.23)	(0.26)	(0.27)	(0.34)	(0.40)	(0.32)	(0.40)	(0.40)	(0.36)	(0.36)	(0.41)
Other religion	-0.26**	0.33^{+}	0.05	0.17	0.05	-0.36	-0.31	0.26	0.46	-0.38	0.81*	-0.35
	(0.10)	(0.18)	(0.25)	(0.23)	(0.23)	(0.37)	(0.35)	(0.33)	(0.47)	(0.26)	(0.40)	(0.54)
Atheist family	0.16	-0.11	-0.63	0.04	-0.53	0.29	-0.57	-0.21	0.86^{+}	-1.05*	2.75**	
	(0.32)	(0.31)	(0.57)	(0.38)	(0.61)	(0.58)	(0.49)	(0.56)	(0.52)	(0.52)	(0.53)	
AFQT score	0.49**	-0.45*	-0.31*	0.23	-0.86**	0.55^{+}	0.61**	-0.66*	-0.11	0.71**	-0.89**	-0.16
below the mean	(0.15)	(0.18)	(0.15)	(0.16)	(0.17)	(0.30)	(0.21)	(0.27)	(0.35)	(0.15)	(0.22)	(0.13)
Mother's	-0.10**	0.14**	0.00	-0.03	0.02	0.00	-0.04^{+}	0.18**	-0.07*	-0.09**	0.12**	0.04
education	(0.03)	(0.04)	(0.03)	(0.04)	(0.05)	(0.07)	(0.02)	(0.05)	(0.03)	(0.02)	(0.03)	(0.04)
Two-parent	-0.21	0.19	0.10	-0.37*	0.35	0.26	-0.19	0.20	0.10	0.22	0.19	-0.58**
house at 14	(0.13)	(0.15)	(0.20)	(0.18)	(0.26)	(0.25)	(0.16)	(0.19)	(0.18)	(0.31)	(0.43)	(0.12)
Currently	0.70**	-1.85**	0.56**	0.63**	-2.08**	0.19	-0.15	-0.05	0.34	0.58^{+}	-1.87**	0.25
married	(0.20)	(0.28)	(0.17)	(0.21)	(0.45)	(0.23)	(0.29)	(0.44)	(0.34)	(0.30)	(0.34)	(0.33)
Contrac. before	-0.40**	0.43**	0.13	-0.54**	0.88**	0.16	-0.06	0.31	-0.26	-0.76**	1.12**	0.14
1st pregnancy	(0.11)	(0.15)	(0.16)	(0.13)	(0.21)	(0.18)	(0.15)	(0.20)	(0.17)	(0.13)	(0.16)	(0.20)
Constant	5.14**	-5.36**	-5.94**	9.08**	-12.81**	-10.28**	11.39**	-13.38**	-10.02**	3.43**	-4.56**	-3.14**
	(0.48)	(0.64)	(0.48)	(0.78)	(1.19)	(1.41)	(0.96)	(0.97)	(0.64)	(0.62)	(0.70)	(0.33)
E(sample)	595	564	528	398	327	284	570	428	522	324	296	287

Note: **significant at 1%; * at 5%; $^{+}$ at 10%. Standard errors clustered by state. The sample is limited to women who are pregnant for the first time. **Column B**: dependent variable equals 1 if 1st pregnancy ended in live birth, 0 otherwise. **Column A**: equals 1 if 1st pregnancy was terminated, 0 otherwise. **Column M**: equals 1 if 1st pregnancy ended in miscarriage, 0 otherwise. All models include state and calendar year fixed effects and a cubic polynomial for age; dummies for calendar months are excluded. Excluded category for religion is women raised in Catholic families.

Appendix D (continued)

Table D-4: Probit coefficient estimates for models with a "legally eligible" indicator, women age 18-20

	Non-poor White				Poor White				Black			
	P	**B**	**A**	**M**	**P**	**B**	**A**	**M**	**P**	**B**	**A**	**M**
Eligible to drink legally	-0.15* (0.06)	1.14** (0.31)	-1.32* (0.60)	-0.86* (0.37)	0.21+ (0.11)	-0.23 (0.72)	2.05 (1.90)	1.11 (0.81)	-0.11+ (0.06)	-0.2 (0.37)	1.02* (0.49)	-0.01 (0.43)
Baptist family	-0.02 (0.07)	0.25 (0.20)	0.06 (0.29)	-0.32 (0.29)	0.18+ (0.10)	1.48* (0.58)	0.86 (1.76)	-2.63** (1.01)	-0.02 (0.10)	0.06 (0.43)	-0.36 (0.67)	0.25 (0.66)
Other religion	-0.04 (0.04)	-0.21 (0.19)	0.43* (0.17)	-0.02 (0.27)	-0.01 (0.11)	0.73* (0.36)	0.58 (0.97)	-1.56 (1.14)	-0.04 (0.11)	-0.26 (0.42)	-0.08 (0.44)	0.75 (0.76)
Atheist family	0.02 (0.10)	-0.12 (0.46)	0.88+ (0.47)	-0.29 (0.20)	0.39** (0.13)	1.89+ (1.08)	2.15* (1.03)	2.36* (1.08)	0.04 (0.18)	0.09 (0.69)	-0.76 (0.69)	-0.26 (0.35)
AFQT score below the mean	0.09+ (0.05)	0.31 (0.19)	-0.13 (0.27)	0.04 (0.04)	0.16* (0.08)	0.29 (0.29)	-3.61** (0.75)	-0.22 (0.16)	0.18* (0.09)	0.79** (0.27)	-0.88* (0.37)	-0.15** (0.05)
Mother's education	-0.04** (0.01)	-0.16** (0.05)	0.22** (0.06)	0.04 (0.25)	0.00 (0.01)	0.11 (0.09)	-0.49** (0.16)	1.31+ (0.75)	-0.02 (0.01)	0.09* (0.04)	0.03 (0.09)	0.17 (0.28)
Two-parent house at 14	-0.12** (0.04)	0.00 (0.16)	-0.05 (0.17)	0.57** (0.21)	-0.04 (0.07)	-0.82* (0.37)	-0.02 (0.76)	2.22** (0.66)	-0.16** (0.06)	-0.29 (0.19)	0.38 (0.27)	0.41 (0.46)
Currently married	0.63** (0.05)	0.83** (0.25)	-2.25** (0.52)	0.17 (0.20)	0.66** (0.08)	0.39 (0.36)		-1.60 (0.98)	0.56** (0.11)	-0.14 (0.40)	-0.09 (0.54)	-0.26 (0.21)
Contrac. before 1st pregnancy	-0.01 (0.04)	-0.48* (0.20)	0.58* (0.29)	0.64 (0.86)	-0.05 (0.07)	-0.59+ (0.34)	3.24* (1.50)	2.54 (3.24)	0.03 (0.06)	-0.24 (0.19)	0.93** (0.31)	-3.17* (1.41)
Constant	-6.85* (3.01)	12.36 (13.65)	-13.72 (18.94)	-13.07 (15.58)	(5.52)	8.17 (24.35)	-229.3** (76.38)	-48.24 (56.33)	-6.30 (4.04)	-33.41 (25.47)	34.18 (35.98)	50.67* (24.79)
E(sample)	28,994	359	318	327	12,269	157	79	108	14,450	247	148	210

Note: ** significant at 1%; * at 5%; + at 10%. Standard errors clustered by state. All models include state and calendar year fixed effects and a cubic polynomial for age. In addition, models in columns P include calendar month fixed effects. Excluded category for religion is women raised in Catholic families. **Column P**: dependent variable equals 1 if pregnant for the first time, 0 otherwise. **Column B (A, M)**: equals 1 if 1st pregnancy ended in birth (abortion, miscarriage), 0 otherwise; the sample is limited to pregnant women. The estimation samples are smaller than the initial samples due to collinearity.