Communication

The Effects of Minimum Legal Drinking Ages on Teen Childbearing

Thomas S. Dee

ABSTRACT

This study provides empirical evidence on the structural relationship between alcohol use and teen childbearing by exploiting the exogenous variation in youth alcohol availability generated by changes in state minimum legal drinking ages. The reduced-form childbearing models are based on state-level panel data and two-way fixed effect specifications as well as models that incorporate as controls the contemporaneous childbearing data from older women who were unaffected by the state changes in youth alcohol policy. The results indicate that alcohol availability and use have large, independent, and statistically significant effects on childbearing among black teens but not necessarily among white teens.

I. Introduction

In the United States, teen childbearing is widely viewed as a major social problem. This concern has been motivated by a number of factors including the broad perception that teen childbearing has dire educational and economic consequences for teen mothers as well as adverse consequences for their children. This concern has also been motivated by the high rates of teen childbearing in the United

Thomas S. Dee is a professor of economics at Swarthmore College and a Faculty Research Fellow at the National Bureau of Economic Research. *He would like to thank William Evans, David Ribar, and two anonymous referees for helpful comments. The usual caveats apply. The data used in this article can be obtained beginning May 2002 through April 2005 from the author, Swarthmore College, Swarthmore, PA 19081, tdee1@swarthmore.edu.*

States relative to other industrialized nations, the increases in teen childbearing in the mid-1980s, the prevalence of sexually transmitted diseases, and the frequently coercive nature of early sexual intercourse (Moore et al. 1998). Because of the extensive social costs associated with teen childbearing, a number of empirical studies have investigated its key determinants.[1]

Several studies have emphasized the empirically large and statistically significant relationships between youth alcohol use and outcomes like risky sexual activity and childbearing.[2] These studies raise the intriguing possibility that teen drinking has a substantial influence on the prevalence of teen childbearing through its effects on sexual frequency and contraceptive use.[3] Therefore, policies that limit youth alcohol consumption may have the side effect of reducing teen childbearing. This study provides direct evidence on this relationship by exploiting the sharp and plausibly exogenous variation in teen alcohol availability generated by changes in state minimum legal drinking ages (MLDA). More specifically, this study evaluates how the widespread increases in state-level MLDA over the 1977–92 period (Table 1) influenced the prevalence of childbearing among black and white teens. Some of the results presented here are based on two-way fixed effect specifications that exploit the panel nature of annual state-level data on teen childbearing rates.[4] The results of these models suggest that teen exposure to a low MLDA (that is, easier access to alcohol) generated statistically significant increases in teen childbearing for both black and white teens.

However, this study also examines these findings further by evaluating additional specifications that include contemporaneous childbearing data from older women who were unaffected by the MLDA changes. This regression framework, which is akin to "differences-in-differences-in-differences" (DDD) estimation, relies on the childbearing rates among older women as controls for the unobserved shocks specific to each state and year combination (Gruber 1994). The results of these models suggest that the links between MLDA exposure and teen childbearing are statistically insignificant for white teens.[5] However, these models also indicate that the estimated links between MLDA exposure and teen childbearing persist among black teens. More specifically, these results consistently indicate that the movement to an MLDA of 21 reduced childbearing rates among black teens by roughly 6 percent. Overall,

1. Maynard (1997) estimates that the annual social costs of teen childbearing are roughly $15 billion. Interestingly, this cost figure consists largely of the costs to fathers and children. For overviews of research on the determinants of teen childbearing see the several contributions in Maynard (1997), Lundberg and Plotnick (1995), and Hofferth and Hayes (1987).

2. The author's calculations based on data from the National Education Longitudinal Study of 1988 (NELS-88) suggest that teens who binge-drink are, on average, as much as 63 percent more likely to become teen mothers. For other evidence on the relationship between alcohol use, youth sexual activity, and childbearing see Shrier et al. (1996), Donovan and McEwan (1995), Leigh and Stall (1993), and Mensch and Kandel (1992).

3. The structural relationships between alcohol use and teen childbearing may be substantially more complex than a simple model that relates a female's own drinking and childbearing. This study presents reduced-form evaluations that inform whether alcohol use and availability have a causal effect on teen childbearing. However, this evidence can only indirectly address the underlying structural relationships.

4. This identification strategy is validated by counterfactual estimations based on the birth rate among older women (aged 25–29) who were plausibly unaffected by the MLDA changes.

5. But the implications of this fragility are not entirely self-evident since the DDD estimation framework leaves limited sample variation in white teen childbearing.

Table 1
Effective State Minimum Legal Drinking Ages, 1977–92

	Number of States by Minimum Legal Drinking Age			
Birth Year	18	19	20	21
1977	30	7	1	13
1978	30	8	1	12
1979	29	8	2	12
1980	23	11	4	13
1981	20	13	4	14
1982	18	13	6	14
1983	16	15	6	14
1984	11	18	7	15
1985	7	19	6	19
1986	7	15	8	21
1987	6	9	3	33
1988	0	5	4	42
1989	0	0	3	48
1990	0	0	0	51
1991	0	0	0	51
1992	0	0	0	51

This classification is based on the MLDA in effect for the largest share of the period during which a given year's full-term births were conceived.

the results of this study provide important evidence that alcohol use and availability are causal and empirically relevant determinants of teen childbearing, particularly for black teens. This implies that effective restrictions on youth alcohol availability and use can lead to sizable reductions in the prevalence of teen childbearing. The persistence of these effects among black teens may also point to the interactive effects of other important contextual factors such as their lower rate of contraceptive use and the relatively small age gap between black teen mothers and the fathers of their children. Although any such interpretation of the indirect evidence is highly qualified, it does underscore a useful direction for further research on what appears to be a key determinant of teen childbearing.

II. Teen Childbearing and Alcohol

Alcohol use and availability may influence teen childbearing in a number of ways. One of the most straightforward formulations of the possible, underlying structural relationship is that alcohol use by young females promotes childbearing through disinhibition and impaired judgement. There is corresponding evidence, which indicates that female teens who drink are more sexually active and less likely to use contraceptives. For example, Shrier et al. (1996) present evidence that teens who drink have more sexual partners and are less likely to use condoms. Similarly, prior evidence indicates that young white females who drink alcohol are more likely to become teen mothers (Mensch and Kendel 1992). However, not all studies have concluded that youth alcohol use and risky sexual activity actually are significantly related (for example, Senf and Price 1994; Fortenberry et al. 1997). And the results of studies that do link alcohol use and sexual activity may be misleading if these correlations reflect unobserved individual heterogeneity instead of underlying causal mechanisms. Furthermore, it should be noted that alcohol use among males could also play an important role, especially given the sometimes coercive nature of early sexual activity. Unfortunately, there is surprisingly little prior evidence regarding the fathers of children born to teen mothers. However, there are at least two stylized facts worth emphasizing. These fathers tend to drink more heavily than males who do not father children (Spingarn and DuRant 1996; Larson et al. 1996). And the fathers are, on average, older than the teen mothers. For example, in a recent study of school-age mothers (Larson et al. 1996), nearly 40 percent of the fathers were aged 20 years or more. Similarly, using data from the National Longitudinal Survey of Youth (NLSY), Brien and Willis (1997) find that among nonblack mothers aged 17–19, nearly 58 percent of the fathers were older than 22 years old. However, for black teens, this age gap was much smaller: among mothers aged 17–19, only about 42 percent of fathers were over 22 years old. The influence of males may be particularly relevant in this context since the focus of this study's evaluations is on youth-specific alcohol policies. The joint age distributions of teen mothers and their partners suggest that such policies may not be entirely constraining, particularly among whites.

The identification strategy employed in this study is based on the extensive state-level experiences with increasing minimum legal drinking ages (MLDA) over the 1977–92 period (See Table 1). The robust effects of MLDA on teen drinking are well documented in studies of teen alcohol use as well as studies of related outcomes like traffic fatalities.[6] The available evidence clearly indicates that the within-state movement to higher MLDA has generated empirically large decreases in alcohol use among young adults. For example, Dee (1999) finds that exposure to an MLDA of 18 increased the prevalence of heavy or binge drinking among teens by 3.1 percentage points, an 8.4 percent increase relative to the mean prevalence of this drinking behavior in the sample. Those results were based on pooled cross-sections from the annual Monitoring the Future (MTF) surveys over the 1977–92 period when

6. However, this study explicitly eschews the frequent use of beer taxes as an identification strategy. Several past studies have concluded that teen drinking is highly responsive to beer taxes (Grossman, Coate and Arluck, 1987; Grossman et al. 1994; Coate and Grossman, 1988; Kenkel, 1993; Cook and Moore, 1994). But the identification strategies in those studies relied exclusively on the cross-state variation in beer taxes. More recent studies (for example, Dee 1999, Mast, Benson and Rasmussen 1999, DiNardo and Lemiuex 1996, Dee and Evans, forthcoming) which examine within-state changes suggest that beer taxes actually have relatively small and statistically insignificant effects on teen drinking. This sensitivity to state fixed effects is not entirely definitive since it removes much of the limited sample variation in beer taxes. However, even in such models, there is sufficient power to reject the conventional tax estimates based on cross-state identifications (Dee 1999).

state-level MLDA were being increased. One caveat that should be noted with regard to that evidence is that, since all of the MTF respondents were high school seniors, the policy responsiveness of their alcohol use may not generalize well to those on the margins of becoming a teen mother. However, the evidence from reduced-form models of traffic fatalities does indicate that the within-state variation in MLDA influenced alcohol availability and use for the entire population of young adults. Furthermore, there is evidence that the within-state movement to higher MLDA was independently given, in part because of the strong Federal compulsion over this period to restrict youth alcohol availability.[7]

III. Data and Specifications

Reduced-form models that relate the movement to higher MLDA with the variation in teen childbearing can provide novel and policy-relevant evidence on the welfare consequences associated with youth alcohol availability. In particular, by exploiting the sharp and plausibly exogenous within-state variation in alcohol availability, such models can generate more definitive evidence on whether alcohol use and availability have causal effects on teen childbearing. The childbearing data for these evaluations were drawn from various editions of the *Vital Statistics of the United States* for each state and the District of Columbia. More specifically, I constructed panel data on childbearing rates by state, year, age, and race for the 1977–92 period. The population data that convert the data on live births into rates are the mid-year population estimates generated by the U.S. Census Bureau for each state, year, age, and racial group. The state, year, and race-specific data on births and population are consistently available over this period only for five-year age groups. Childbearing rates were constructed by state, year, and race for three of the five-year age groups: 15–19 year olds, 20–24 year olds, and 25–29 year olds (appendix table).[8] As in much of the prior literature, one of the specifications employed here for analyzing these panel data is a two-way fixed effects model where the dependent variable, Y_{st}, is the natural log of the birth rate per 1,000 in the state-year population under study. More specifically, this equation takes the following basic form:

$$(1) \quad Y_{st} = X_{st}\beta + M_{st}\gamma + u_s + v_t + \varepsilon_{st}$$

where X_{st} represents the observed determinants of teen childbearing in state s during year t and ε_{st} is a mean-zero error term. The terms, u_s and v_t, are state and year fixed effects. These controls exploit the panel nature of the available data by providing unambiguous controls for the unobserved childbearing determinants specific to each

7. Dee and Evans (forthcoming) discuss other anecdotal and ad-hoc empirical evidence in support of this conventional assumption.

8. The MLDA variation is relevant for only some of the births among 20–24 year olds for whom conception occurred at ages ranging from 19 to 24. It is presumably irrelevant for births among 25–29 year olds. However, as is discussed shortly, these data provide the basis for a compelling validation of the estimation approach as well as possible controls for state and year-specific birth shocks. This data set has 816 observations on childbearing rates (51 states by 16 years) for each age and race. However, models on childbearing among blacks exclude Vermont, which had no births to black mothers in some years. The results presented here are similar in linear probability models that include these data.

state and year. The term, M_{st}, represents the salient MLDA characteristics in a given state and year. The MLDA controls employed here represent the variation between minimum ages of 18, 19, 20, and 21 in an unrestrictive manner. These controls consist of one to three binary indicators and were constructed by identifying the MLDA policy in effect for the largest part of the year during which a calendar year's births were ostensibly conceived.[9] The models evaluated here also introduce two state macroeconomic variables (that is, state unemployment rate and real per capita personal income defined for the year prior to the one in which the births occurred) as well as three additional regressors representing other important state policies. One state policy variable is a measure of welfare generosity in the year prior to the observed births: real maximum AFDC benefit level for a family of three. Moffit (1992) discusses the evidence linking such welfare incentives to teen childbearing. The other two state policy measures reflect abortion access in the year prior to the observed teen births. One is a binary indicator for the presence of an enforced parental consent or notification law. The second variable is a binary indicator that represents the presence of enforced restrictions on the Medicaid funding of abortions.[10] There is some evidence that such policies influenced the number of abortions (for example, Blank, George, and London 1996). However, Kane and Staiger (1996) concluded that these restrictive abortion policies had weak (but possibly negative) effects on the prevalence of teen childbearing.[11]

There are undoubtedly other attitudinal, institutional, and economic variables that influenced the within-state variation in teen childbearing but are difficult to measure and are necessarily omitted from these specifications. The plausibly exogenous nature of the within-state MLDA variation suggests that such omitted variables are unlikely to impart a problematic bias. However, a compelling, ad hoc check of this important specification concern can be conducted by exploiting the panel data on childbearing rates among 25–29 year olds. The logic of such counterfactual estimations is straightforward. If the two-way fixed effects models presented here are generating reliable inferences about the true relationship between MLDA changes and teen childbearing, then we should also observe, in similarly specified models of childbearing among older females, small and statistically insignificant effects associated with the MLDA variables. In contrast, if there were significant links between the MLDA changes and childbearing among older women, it would suggest the existence of important and overlooked specification errors (for example, omitted determinants of childbearing that are confounded with the MLDA variation). Table 2 presents evidence from such evaluations. More specifically, Table 2 reports the estimated effect of an MLDA of 21 on childbearing rates among white and black 25–29 year olds. Given the likely presence of heteroscedasticity, the inferences are

9. These variables incorporate the grandfathering of some changes (O'Malley and Wagenaar 1991) to identify the effective MLDA in each state and year.

10. I would like to thank David Ribar for generously providing data on Medicaid restrictions. The variable representing parental consent and notification laws was defined using the information available in Haas-Wilson (1996, Table 1), the Alan Guttmacher Institute's "The Status of Major Abortion-Related Laws in the States," and state-specific Lexis-Nexis searches.

11. In particular, these policies did not appear to have an effect in models that conditioned on state-specific trends in childbearing. Joyce and Kaestner (1996) also present evidence that such laws had fairly small effects on pregnancy resolution.

Table 2
The Estimated Effect of a Minimum Legal Drinking Age of 21 on Childbearing among 25–29 year old Whites, Weighted and Unweighted Results

Estimation Method	Model (1)	Model (2)	Model (3)
White women			
Ordinary least squares	−0.0005	−0.004	−0.008
	(0.013)	(0.012)	(0.011)
R^2	0.8690	0.8728	0.8745
Weighted least squares	−0.025***	−0.024***	−0.024***
	(0.008)	(0.008)	(0.008)
R^2	0.7776	0.7862	0.7974
Black women			
Ordinary least squares	0.021	0.003	0.002
	(0.021)	(0.020)	(0.020)
R^2	0.6539	0.6665	0.6692
Weighted least squares	0.002	−0.002	0.002
	(0.008)	(0.008)	(0.007)
R^2	0.7111	0.7633	0.7839
Macroeconomic variables	No	Yes	Yes
State policy variables	No	No	Yes

All models include state and year fixed effects. The dependent variable is the natural log of the childbearing rate per 1,000 25–29 year olds. The weights in the WLS models are the state-year population estimates. Heteroscedastic-consistent standard errors are reported in parentheses for the OLS models.
*** Statistically significant with a p-value $< .01$.

based on heteroscedastic-consistent standard errors. These models generate the plausible results that exposure to an MLDA of 21 had small and statistically insignificant effects on childbearing rates among these older women.

These counterfactuals can also provide some guidance as to the treatment of heteroscedasticity. Several related studies have emphasized the likely heteroscedasticity in grouped vital statistics data like these and adopted a generalized least square (GLS) approach based on using the estimated population size as a weight (for example, Kane and Staiger 1996, Haas-Wilson 1996). The counterfactual results presented in the bottom panel of Table 2 are similarly based on population-weighted least squares. Implausibly, these WLS results suggest that the movement to an MLDA of 21 led to fairly large and statistically significant reductions in childbearing rates among white 25–29 year olds. The model results in the bottom panel of Table 2 are based on childbearing among black women and do not exhibit sensitivity to the use of weighted least squares. But the implausibility of the WLS results in the models for white women suggests that the population estimates may be inappropriate weights. In light of this concern about the exact form of the underlying heteroscedasticity, this study reports only heteroscedastic-consistent standard errors based on unrestrictive corrections (White 1980). However, the power of these counterfactual estimations should not be overdrawn. In particular, the possibility that these implausible WLS results reflect some other unknown model misspecification cannot be ruled out.[12]

This study also presents evaluation results that are based on a regression model that explicitly relies on childbearing data among older women to provide possibly better controls for unobserved state and year-specific childbearing shocks. More specifically, this second specification takes the following basic form:

$$(2)\quad Y_{ist} = X_{ist}\beta + (M_{st} \times w_i)\gamma + u_s + v_t + w_i + (u_s \times w_i) + (v_t \times w_i) + (u_s \times v_t) + \varepsilon_{ist}.$$

The term, w_i, represents fixed effects for each age group. This unrestrictive model includes a full set of interactions between the age, year and state fixed effects. Identification in this "differences-in-differences-in-differences" (DDD) specification relies effectively on comparing the change in the gap between teen and adult childbearing rates in states that did and did not raise their MLDA (see, for example, Gruber 1994). The interactions between the MLDA variables, M_{st}, and the age fixed effects, w_i, are set to zero for the 25–29 year olds for whom these regulations were presumably irrelevant. The MLDA variables are defined in an unrestrictive manner for the 20–24 year old age group since, at the time of conception, mothers in this age group were as young as 19.[13] Because these models include interactions of state and year fixed effects (that is, $u_s \times v_t$), the variable, X_{ist}, only includes the binary indicator for the parental consent or notification law since the other state-year controls are perfectly collinear. It should be noted that this "DDD" approach assumes that older and younger women share the same childbearing shocks in a given state and year. The two-way fixed effect models, which instead use as controls the within-state childbearing shocks among teens in other states, may actually be preferable. Since there is little basis for distinguishing these approaches a priori, these models are probably best viewed as complementary approaches for exploring the validity of this study's key results.

IV. Results

The results in Tables 3 and 4 are based only on the two-way fixed effect specification and data from teen mothers aged 15–19. Table 3 presents the key results for white mothers. The estimates from the sparsest specification indicate that the within-state variation in MLDA had small and statistically insignificant or implausibly signed effects on white teen childbearing. Similarly, in models that introduce the state-year macroeconomic controls and policy variables, this evidence suggests that the movement to an MLDA of 21 had small and statistically insignificant effects on childbearing rates among white teens. However, the models that include these controls also indicate that exposure to an MLDA of 18 was associated with statistically significant 2.9 to 3.7 percent increases in childbearing rates among white

12. However, this study's main results (that is, MLDA effects for childbearing among black teens) are similar across the two main specifications employed here regardless of the estimation procedure.

13. This implies that an MLDA of 18 was probably irrelevant for this group. However, that hypothesis is tested directly here.

Table 3
Teen Childbearing and Alcohol Availability, 15–19 Year Old Whites, 1977–92

Variable	Model (1)	Model (2)	Model (3)	Model (4)	Model (5)	Model (6)
MLDA of 18	—	−0.0005	—	0.029**	—	0.037***
		(0.016)		(0.014)		(0.014)
MLDA of 19	—	−0.011	—	−0.004	—	−0.0001
		(0.015)		(0.014)		(0.014)
MLDA of 20	—	−0.029**	—	−0.004	—	0.001
		(0.014)		(0.014)		(0.014)
MLDA of 21	0.013	—	−0.006	—	−0.012	—
	(0.013)		(0.012)		(0.012)	
R^2	0.9354	0.9357	0.9483	0.9489	0.9509	0.9517
Macroeconomic variables	No	No	Yes	Yes	Yes	Yes
State policy variables	No	No	No	No	Yes	Yes

All models include state and year fixed effects. The dependent variable is the natural log of the childbearing rate per 1,000 15–19 year olds. Heteroscedastic-consistent standard errors are reported in parentheses.
** Statistically significant with a p-value $< .05$.
*** Statistically significant with a p-value $< .01$.

teens (Models 4 and 6). These results suggest that alcohol availability and, by implication, alcohol use have a causal impact on childbearing among white teens. Similar to the evidence from models of teen alcohol use (Dee 1999), this particular heterogeneity indicates that the relevant change in alcohol availability was the movement away from an MLDA of 18 as opposed to the movement to an MLDA of 21. This pattern of results also provides some indirect evidence that for white teen mothers, the alcohol availability of the older males who typically father their children was not distinctly relevant.

The results for two-way fixed effect models of childbearing among black teens are presented in Table 4. These estimates also suggest that the within-state variation in MLDA exposure had large and statistically significant effects on childbearing rates among black teens. For example, in the models that include the macroeconomic and policy variables, these results indicate that the movement to an MLDA of 21 reduced the prevalence of teen childbearing by a statistically significant 6.1 to 6.4 percent (Models 3 and 5). Interestingly, these results also suggest that these effects were largely driven by teen exposure to an MLDA of 19 and not by exposure to an MLDA of 18, which had a positive but relatively small and insignificant effect. In particular, for those exposed to an MLDA of 19, the estimated increase in childbearing was 8.2 to 8.9 percent. In light of the statistical precision, however, the difference between the estimated effects of an MLDA of 18 and an MLDA of 19 is not very large: the 95 percent confidence interval for the effect of an MLDA of 19 includes the estimated effect of an MLDA of 18.

The results in Tables 3 and 4 suggest that the within-state variation in alcohol

Table 4
Teen Childbearing and Alcohol Availability, 15–19 Year Old Blacks, 1977–92

Variable	Model (1)	Model (2)	Model (3)	Model (4)	Model (5)	Model (6)
MLDA of 18	—	−0.002	—	0.035	—	0.036
		(0.029)		(0.029)		(0.028)
MLDA of 19	—	0.071**	—	0.082***	—	0.089***
		(0.031)		(0.030)		(0.030)
MLDA of 20	—	0.017	—	0.056*	—	0.056*
		(0.030)		(0.029)		(0.029)
MLDA of 21	−0.033	—	−0.061***	—	−0.064***	—
	(0.024)		(0.024)		(0.023)	
R^2	0.5460	0.5503	0.5684	0.5701	0.5927	0.5948
Macroeconomic variables	No	No	Yes	Yes	Yes	Yes
State policy variables	No	No	No	No	Yes	Yes

All models include state and year fixed effects. The dependent variable is the natural log of the childbearing rate per 1,000 15–19 year olds. Heteroscedastic-consistent standard errors are reported in parentheses.
* Statistically significant with a p-value < .10.
** Statistically significant with a p-value < .05.
*** Statistically significant with a p-value < .01.

variation created by changes in minimum legal drinking ages had significant effects on childbearing among both black and white teens. However, the results from such two-way fixed effect models may be misleading if there were other reasons that the within-state variation in teen childbearing over this period varied with the state-specific MLDA changes. Table 5 provides evidence on the possible presence of such an omitted variable bias by reporting the results from DDD specifications that include childbearing data from white mothers aged 15–19, 20–24, and 25–29. Unlike the results in Table 3, the results in the first two columns of Table 5 suggest that the estimated effects of MLDA exposure on childbearing among young white women were small and statistically insignificant or implausibly signed. However, the implications of this sensitivity are not entirely clear. Because the within-state variation in white childbearing rates becomes limited in the "DDD" models that include state-by-year fixed effects, the true effects of MLDA exposure on white teen childbearing may be best viewed as uncertain. In particular, it should be noted that, in Model 2 of Table 5, the estimated effect of exposure to an MLDA of 18 on white teen childbearing includes in its 95 percent confidence interval the largest estimate from Table 3. Furthermore, in shorter panels that focus only on the earliest MLDA changes (for example, 1977–82), there are statistically significant MLDA effects similar to those in Table 3 in both specifications.

The results of the "DDD" specifications for childbearing among black women are largely consistent with those based on the two-way fixed effect specification. The results in Table 5 indicate that the movement to an MLDA of 21 reduced

Table 5
Teen Childbearing and Alcohol Availability, 15–19, 20–24 and 25–29 Year Olds, 1977–92

	White		Black	
Variable	Model (1)	Model (2)	Model (3)	Model (4)
MLDA of 18 × Age 15–19	—	−0.006 (0.022)	—	0.036* (0.021)
MLDA of 19 × Age 15–19	—	−0.006 (0.015)	—	0.088*** (0.025)
MLDA of 20 × Age 15–19	—	−0.029** (0.014)	—	0.030 (0.031)
MLDA of 21 × Age 15–19	0.012 (0.015)	—	−0.055*** (0.018)	—
MLDA of 18 × Age 20–24	—	0.025 (0.023)	—	0.034 (0.021)
MLDA of 19 × Age 20–24	—	0.011 (0.014)	—	0.064*** (0.022)
MLDA of 20 × Age 20–24	—	0.008 (0.013)	—	0.047 (0.030)
MLDA of 21 × Age 20–24	−0.015 (0.015)	—	−0.050*** (0.018)	—
R^2	0.9898	0.9898	0.9258	0.9261

All models include a binary indicator for a parental consent/notification law, state, year, and age fixed effects and interactions between state and age fixed effects, year and age fixed effects, and state and year fixed effects. The dependent variable is the natural log of the childbearing rate per 1,000 in the given age group. Heteroscedastic-consistent standard errors are reported in parentheses.
* Statistically significant with a p-value < .10.
** Statistically significant with a p-value < .05.
*** Statistically significant with a p-value < .01.

childbearing rates among black teens by an estimated 5.5 percent. The estimated effects are again quite large for teen exposure to an MLDA of 19, which increased childbearing rates by an estimated 8.8 percent. Exposure to an MLDA of 18 was also associated with a relatively small (3.6 percent) but weakly significant increase in childbearing among black teens. The relatively small and often statistically insignificant effects of exposure to an MLDA of 20 are not entirely surprising since the sample variation in an MLDA of 20 was limited to relatively few states and time periods (Table 1). The estimated effects of MLDA exposure on childbearing among 20–24 year old blacks deserve special mention. Perhaps surprisingly, these results indicate that the movement to an MLDA of 21 reduced childbearing among this group by 5 percent. These results are not necessarily as anomalous as they might initially seem. A large share of the conceptions in this age group occurred at the age of 19 and 20, when MLDA beneath 21 may have constrained alcohol availability. Furthermore, as would be expected, these estimated effects among 20–24 year olds are smaller than for teens. Similarly, the estimated effects of an MLDA of 18 are smaller and statistically insignificant for this group, which is to be expected since that MLDA was particularly irrelevant for this age group. Nonetheless, I also reestimated these models excluding data childbearing data from the 20–24 year olds and found that the key MLDA results for teens were quite similar.

Overall, these evaluation results indicate that alcohol availability and, by implication, alcohol use are significant determinants of teen childbearing. However, the "DDD" models suggested that the MLDA variation only influenced childbearing among black teens. Three possible and not mutually exclusive explanations for this race-specific heterogeneity are worth underscoring. First, as suggested earlier, these results may simply reflect the weak statistical power implied by the relatively limited within-state variation in teen childbearing among whites. The relatively high R^2 and large standard errors from models of white childbearing (Table 5) are suggestive of this possibility. However, the pattern to the MLDA results in Tables 3 through 5 may also reflect race-specific structural responses and the unique, race-specific context in which alcohol use and availability influence teen childbearing. While it is admittedly speculative to rely on the pattern of reduced-form results to draw strong inferences about the nature of the underlying structural relationships, there are at least two suggestive possibilities. For example, there is evidence that black teens are substantially more likely to be sexually active and less likely to use contraception (Levine 2001). These patterns of behavior are likely to interact with youth alcohol use and availability in a manner that amplifies the effects of MLDA among black teens. Furthermore, while the fathers of the children born to teen mothers, are, on average, several years older than the mothers, there is evidence that this gap is much smaller for black teens (Brien and Willis 1997). This stylized fact suggests that youth-specific constraints on alcohol availability would be particularly relevant for childbearing among black teens.

V. Conclusions

The widely held view that teen childbearing is one of the United States' major social problems has focused considerable attention on its key determinants. In particular, several recent studies of the determinants of teen childbearing have noted the strong links between alcohol use, sexual activity and childbearing among youths. Teens who drink heavily are as much as 63 percent more likely to be teen mothers. The magnitude of this correlation suggests that teen childbearing and its attendant social costs may be an important life-cycle consequence of teen drinking. These links also raise the intriguing possibility that reductions in youth alcohol availability can lead to empirically important reductions in the prevalence of teen childbearing. However, the policy relevance of these suggestive correlations hinges critically on whether alcohol use has an independent effect on childbearing outcomes. This study provided evidence on this critical question by exploiting the sharp and plausibly exogenous variation in youth alcohol availability generated by changes in states' MLDA. More specifically, the key evaluations presented here addressed the effects of MLDA increases in reduced-form models of teen childbearing based on state-level panel data from the 1977–92 period.

The results of evaluations based on two-way fixed effect specifications suggested that the nationwide movement to higher MLDA led to large and statistically significant reductions in teen childbearing among both white and black teens. This conventional panel data model was validated by counterfactual estimations, which found that, in similarly specified models, the movement to higher MLDA had small and statistically imprecise effects on the childbearing rates of older women. However, this study also examined the validity of these results by evaluating ''differences-in-differences-in-differences'' models that included childbearing data from older women as controls for the unobserved childbearing shocks specific to each state and year. The results of these models indicated that the estimated MLDA effects were fragile for white teens but persistent among black teens. In particular, these estimates suggested that the nationwide movement to an MLDA of 21 reduced childbearing by roughly 6 percent among black teens (Table 5). The persistent link between alcohol availability and childbearing among black teens is a particularly unusual result since some prior studies have emphasized the unresponsiveness of black teen childbearing to other public policies (for example, Lundberg and Plotnick 1995). Furthermore, these results also provide motivation and some guidance for future research on an important determinant of teen childbearing. In particular, the heterogeneity of these results by race suggests provocatively that the alcohol availability of the fathers as well as related sexual behaviors (for example, sexual frequency and contraceptive use) interact to determine the effects of alcohol availability and use on teen childbearing.

Table A1
Descriptive Statistics, State-Level Panel Data, 1977–92 (Standard Deviations in Parentheses)

Variable	Mean
Births per 1,000 white women	
Age 15–19	44.4
	(13.1)
Age 20–24	108.9
	(28.2)
Age 25–29	112.6
	(19.6)
Births per 1,000 black women	
Age 15–19	104.2
	(28.2)
Age 20–24	165.8
	(45.8)
Age 25–29	126.3
	(39.3)
MLDA of 18	0.24
	(0.43)
MLDA of 19	0.17
	(0.38)
MLDA of 20	0.07
	(0.25)
Real AFDC maximum benefit for a family of three	337
	(130)
Medicaid funding restriction	0.61
	(0.46)
Parental consent or notification law	0.19
	(0.39)
State unemployment rate	0.07
	(0.02)
Real state personal income per capita	12,256
	(2,157)

References

Blank, R. M., C. C. George, and R. A. London. 1996. ''State Abortion Rates—the Impact of Policies, Providers, Politics, Demographics, and Economic Environment.'' *Journal of Health Economics* 15(5):513–53.

Brien, M. J., and R. J. Willis. 1997. ''Costs and Consequences for the Fathers.'' In *Kids Having Kids: Economic Costs and Social Consequences of Teen Pregnancy*, ed. R. Maynard, Washington D.C.: Urban Institute Press.

Coate, D., and M. Grossman. 1988. ''Effects of Alcoholic Beverage Prices and Legal Drinking Ages on Youth Alcohol Use.'' *Journal of Law and Economics* 31(1):145–71.

Cook, P. J., and M. J. Moore. 1994. ''This Tax's For You: The Case for Higher Beer Taxes.'' *National Tax Journal* 47:559–73.

Dee, T. S. 1999. ''State Alcohol Policies, Teen Drinking and Traffic Fatalities.'' *Journal of Public Economics* 72(2):289–315.

Dee, T. S., and W. N. Evans. ''Teen Drinking and Educational Attainment: Evidence From Two-Sample Instrumental Variables (TSIV) Estimates.'' *Journal of Labor Economics.* Forthcoming.

DiNardo, J., and T. Lemieux. 1996. ''The Effect of State Drinking Age Laws on the Consumption of Alcohol and Marijuana by High School Seniors.'' Mimeo.

Donovan, C., and R. McEwan. 1995. ''A Review of the Literature Examining the Relationship Between Alcohol Use and HIV-Related Sexual Risk-Taking in Young People.'' *Addiction* 90:319–28.

Fortenberry, J. D., D. P. Orr, B. P. Katz, E. J. Brizendine, and and M. J. Blythe. 1997. ''Sex Under the Influence: A Diary Self-Report Study of Substance Use and Sexual Behavior Among Adolescent Women.'' *Sexually Transmitted Diseases* 24(6):313–19.

Grossman, M., D. Coate, and G. M. Arluck. 1987. ''Price Sensitivity of Alcohol Beverages in the United States.'' In *Control Issues in Alcohol Abuse Prevention: Strategies for States and Communities*, ed. H. D. Holder. Greenwich, Conn.: JAI.

Grossman, M., F. J. Chaloupka, H. Saffer, A. Laixuthai. 1994. ''Effects of Alcohol Price Policy on Youth: A Summary of Economic Research.'' *Journal of Research on Adolescence* 4(2):347–64.

Gruber, J. 1994. ''The Incidence of Mandated Maternity Benefits'' *American Economic Review* 84(3):622–41.

Haas-Wilson, D. 1996. ''The Impact of State Abortion Restrictions on Minors' Demand for Abortions.'' *Journal of Human Resources* 31(10):140–58.

Hofferth, S., and C. Hayes. 1987. *Risking the Future: Adolescent Sexuality, Pregnancy and Childbearing*, Vol 2. Washington, D.C.: National Academy Press.

Joyce, T., and R. Kaestner. 1996. ''State Reproductive Policies and Adolescent Pregnancy Resolution: the Case of Parental Involvement Laws.'' *Journal of Health Economics* 15: 579–607.

Kane, T. J., and D. Staiger. 1996. ''Teen Motherhood and Abortion Access.'' *Quarterly Journal of Economics*, 477–506.

Kenkel, D. 1993. ''Drinking, Driving, and Deterrence: The Social Costs of Alternative Policies.'' *Journal of Law and Economics* 36:877–914.

Larson, N. C., J. M. Hussey, M. R. Gillmore, and L. D. Gilchrist. 1996. ''What About Dad? Fathers of Children Born to School-Age Mothers.'' *Families in Society: The Journal of Contemporary Human Services* 77(5):279–89.

Leigh, B. C., and R. Stall. 1993. ''Substance Use and Risky Sexual Behavior for Exposure to HIV: Issues in Methodology, Interpretation, and Prevention.'' *American Psychologist* 48:1035–45.

Levine, P. 2001. ''The Sexual Activity and Birth Control Use of American Teenagers'' In *An Economic Analysis of Risky Behavior Among Youths*, ed. J. Gruber. Cambridge, Mass.: National Bureau of Economic Research.

Lundberg, S., and R. D. Plotnick. 1995. ''Adolescent Premarital Childbearing: Do Economic Incentives Matter?'' *Journal of Labor Economics* 13(2):177–200.

Mast, B. D., B. L. Benson, and D. W. Rasmussen. 1999. ''Beer Taxation and Alcohol-Related Traffic Fatalities.'' *Southern Economic Journal* 66(2):214–49.

Maynard, R. A., ed. 1997. *Kids Having Kids: Economic Costs and Social Consequences of Teen Pregnancy.* Washington, D.C.: Urban Institute Press.

Mensch, B., and D. B. Kandel. 1992.''Drug Use as a Risk Factor for Premarital Teen Pregnancy and Abortion in a National Sample of Young White Women." *Demography* 29(3):409–29.

Moffitt, R. 1992. "Incentive Effects of the U.S. Welfare System—A Review." *Journal of Economic Literature* 30(1):1–61.

Moore, K. A., B. C. Miller, B. W. Sugland, D. R. Morrison, D. A. Glei, and C. Blumenthal. 1998. "Adolescent Sexual Behavior, Pregnancy and Parenthood: A Review of Research and Interventions." Washington, D.C.: Office of the Assistant Secretary for Planning and Evaluation, U.S. Department of Health and Human Services.

O'Malley, P. M., and A. C. Wagenaar. 1991. "Effects of Minimum Drinking Age Laws on Alcohol Use, Related Behaviors and Traffic Crash Involvement Among American Youth: 1976–1987." *Journal of Studies on Alcohol* 52(5):478–91.

Ribar, D. C. 1994. "Teenage Fertility and High School Completion." *Review of Economics and Statistics*, 413–24.

Senf, J. H., and C. Q. Price. 1994. "Young Adults, Alcohol and Condom Use: What is the Connection?" *Journal of Adolescent Health* 15:238–44.

Shrier, L. A., S. J. Emans, E. R. Woods, R. H. DuRant. 1996. "The Association of Sexual Risk Behaviors and Problem Drug Behaviors in High School Students." *Journal of Adolescent Health* 20:377–83.

Spingarn, R. W., and R. H. DuRant. 1996. "Male Adolescents Involved in Pregnancy: Associated Health Risk and Problem Behaviors." *Pediatrics* 98(2):262–68.

White, H. 1980. "A Heteroskedasticity-Consistent Covariance Matrix Estimator and a Direct Test for Heteroskedasticity." *Econometrica* 48:817–38.