Common Law Marriage and Teen Births [1]

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Abstract. Whether Common Law Marriage (CLM) in the US affects teen birth rates is the central question we address. CLM effects were identified through cross-state and time variation, as four states repealed the law over the period of study. Using microdata from Current Population Survey Fertility supplements 1990-2010 and state-level data from CDC Vital Statistics 1988-2012 we found that in the states where CLM was first available but then repealed the odds that teens would become new mothers increased. Births to teens younger than 18 were more responsive to availability of CLM than those to teens aged 18 or 19 or to women in their early twenties. The likelihood of becoming a mother increased where CLM is available in the years prior to its repeal. Teens were more responsive to information about availability of CLM about three years later than to knowing that it is available at the time of potential conception. To the extent that they reduce teen births CLM laws are socially desirable and states that still have CLM may be better off by not repealing the law.

Introduction

In the recent past the US has experienced a dramatic drop in teen birth rates: between 2000 and 2013 the birth rate per 1,000 teens ages 15 to 19 dropped by 44% (Martin, Hamilton and Osterman 2014). However, at 26.5 per thousand the US teen birth rate remains higher than that of other developed countries (United Nations 2013). It is particularly high for black and Hispanic teens (Hamilton et al. 2014). Early parenthood has been associated with lower educational attainment and increased poverty for mothers (Hoffman and Maynard 2008) and lower academic performance and more behavioral problems for children (Martinez, Copen and Abma 2011), including higher rates of teenage pregnancy for daughters (Moore et al. 2014). Consequently, there is great interest in uncovering the factors that contribute to high teen births in the US and a number of programs have been introduced with the goal of reducing these births (Kearney and Levine 2015). This study examined whether Common Law Marriage (CLM), a particular type of state law regulating cohabitation and marriage, affects teen birth rates. No previous studies have attempted to link this law to birth rates.

Previous research found that while declining welfare benefits have a statistically discernible negative impact on teen birth rates their magnitude is small (see Kearney and Levine 2015 and cites therein). Lower marital prospects encouraged non-marital births among young women and teen birth rates dropped with expanded access to state-subsidized family planning and weak labor market conditions (Kearney and Levine 2015 and cites therein). Other factors and policies covered by Kearney and Levine--including abstinence education, mandatory sex education, contraception counseling, state children health insurance policy implementation, and total child support expenditures—did not appear to have an impact on teen birth rates. According to Mechoulan (2011) higher rates of black male incarceration were associated with lower odds of non-marital teenage motherhood among young black women. Some of the factors influencing overall birth rates may also have affected teen births rates, and research has shown that overall birth rates are negatively affected by house prices (Dettling and Kearney 2014), availability of abortions (Levine et al. 1999), and culture (Fernández and Fogli 2009).

Closest to the focus of our research are studies of the effect of divorce laws on overall fertility, and on non-marital fertility in particular. Laws liberalizing divorce have been shown to lead to a reduction in fertility (Stevenson 2007 and Drewianka 2008 in the US; Bellido and Marcen 2014 in Europe). The same studies reported that in both the US and Europe the introduction of unilateral divorce laws led to permanent drops in non-marital fertility (including births to cohabiting couples) while it only led to transitory drops in marital fertility, resulting in an increase in the ratio of marital to non-marital births. Using a sample of close to 20 different legal regimes Ekert-Jaffe and Grossbard (2008) found that the likelihood of an out-of-couple birth was lower when a country or province had community property laws guiding division of assets in case of dissolution. However, such laws had less of an impact on the likelihood that a teen had a child while being out-of-couple.

US states with CLM laws offer their heterosexual residents the option of being considered as married by living together and holding themselves out as spouses by calling each other husband and wife in public, using the same last name, filing joint tax returns, or declaring their marriage on applications, leases, birth certificates and other documents. As of 2016 CLM was effective in 11 states--Alabama, Colorado, Iowa, Kansas, Montana, New Hampshire (posthumously for purposes of inheritance), Oklahoma, Rhode Island, South Carolina, Texas, and Utah--as well as in the Navajo Nation and the District of Columbia. CLM does not require a ceremony or a license. Once established, CLM is like a regular marriage, and it requires an official divorce in case of dissolution. However the marriage start date is often not well defined; marriage can be claimed ex-post unilaterally for the purposes of asset division. Thus it is possible for partners, often husbands, to live in CLM marriage without knowing it because the wife can claim that the CLM existed at the time of divorce. Cohabiting couples who have a child are easily categorized as married in a CLM state. In such a state teenagers who started cohabiting below the legal age of marriage (18 in most states) could possibly claim they are married after they turn 18, especially if they have a child with their partner.[2]

Data availability on CLM is a problem. There is virtually no official data on CLM marriages published by state or local governments, even though some counties encourage residents to register their CLMs. CLM is practiced given that about one hundred legal CLM–related judgments were issued each decade in each state at the federal level for those marriages that are disputed (Lind 2008).

We analyzed data on teen childbearing using data from *CDC Vital Statistics* for the years 1988 to 2012 [3], and the *June Fertility Supplements to the Current Population Surveys[4] (CPS)* for the years 1990, 1992, 1994, 1995, 1998, 2000, 2002, 2004, 2006, 2008, and 2010. The periods covered span the years during which 4 states abolished CLM: Ohio (October 1991), Idaho (1996), Georgia (1997), Pennsylvania (2005).

In previous research we found that CLM laws had a negative impact on couple formation: where and when CLM is available most men and women ages 18 to 35 were less likely to live in couples, married or cohabiting, although this did not hold for all education groups (Grossbard and Vernon 2014). By making cohabitation more costly for the partner with more assets, CLM discouraged cohabitation, and since about half of all marriages start with cohabitation, the law also reduced marriage rates.

Couples often decide to cohabit as a consequence of unplanned pregnancy. In turn, having a child together with someone is one of the grounds that individuals can use to claim marriage unilaterally according to CLM. Therefore it is possible that CLM simultaneously reduces cohabitation and lowers birth rates.

The analyses we present here indicate that the abolition of CLM had a small positive effect on births to teen mothers. In absolute terms, availability of CLM three years in the future had a larger impact than its availability around the time of potential conception. Births to younger

[2] In Nebraska the legal age of marriage is 19. All states require parental consent for minors age 16-18 and minors age 14-15 require permission from Juvenile Court. See www.usmarriagelaws.com/search/united_states/teen_marriage_laws/
[3] http://www.cdc.gov/nchs/products/nvsr.htm and http://www.cdc.gov/nchs/products/mvsr.htm
[4] https://cps.ipums.org/cps/

teens age 15 to 17 were more responsive to the availability of CLM than those to older teens. For women aged 20 to 22 we only found a temporary effect of CLM on births, suggesting that they or their partners did not systematically change their fertility behavior as a function of availability of CLM.

In the next section we expand on the reasons why CLM laws may affect the likelihood that a woman becomes a mother, and how this is likely to vary with teen status, child parity, and ethnicity. We then present the data and the methods, followed by a discussion of the results. We conclude with a summary, some policy implications and suggestions for further research.

CLM Laws, Teen Status and the Likelihood that a Woman is a Mother

In the tradition of the New Home Economics (NHE) pioneered by Gary Becker (1960) and Jacob Mincer (1963) it is assumed that decisions regarding fertility are a function of costs and benefits. However, most economists analyzing fertility assume that the unit making decisions regarding births is a household. In contrast, in line with Ekert-Jaffe and Grossbard (2008), we assume that a woman makes that decision and that her decision is potentially tied to the decision on whether to have a child out-of-couple or after first becoming part of a couple.

A decision-making model

Consider a woman deciding on whether to give birth to a child or not. Her decision is related to other decisions, some of which being more under her control than others. In the US women tend to have control over whether to engage in sex with or without protection, to have an abortion, or to give their child for adoption. They may also decide on whether to have a child alone (a) or in couple. If in couple, they can choose between entering (non-marital) cohabitation (co) or marriage (ma) with a man or a woman. We assume heterosexuality and denote BR as the benefit to cost ratio of a particular way of bringing a child into the world.

For unmarried women living in CLM states, having a child can be a step towards being considered married even without the consent of the baby's father. This may be particularly relevant to pregnant teenagers who may experience social pressure to cohabit with their child's father, which under CLM can effectively mean eventual marriage. The cost of having a child in a CLM state thus includes the expected costs of being unwillingly considered as married. The less the individual considers it beneficial to be married, the higher that cost of having a child in a CLM state. For example, individuals with substantial financial assets, or whose families own such assets, may be particularly reluctant to be placed in a situation that could bring a CLM-forced marriage and are therefore more likely to avoid unprotected sexual relations.

Consider three periods. In Period zero the woman is making the decision regarding having a child in each relationship status that she may enter in Period 1. The expected BR at time zero is

(1) $E(BR) = p_{co}E(BR_{co})+ p_{ma}BR_{ma}+p_{a}BR_{a}$, where

p is the probability of being in each type of relationship in the first period and where $p_{co}+ p_{ma}+ p_{a}= 1$. Men's willingness to have a child and/or to form a couple with her will influence the decision. Having and raising a man's child is a form of Work-In-Household (WiHo) that a mother may supply to men (see Grossbard 2015). The potential benefits of being a mother include resources (monetary, in-kind, and caring time) that men are willing to provide their child's mother and that can be interpreted as payments for WiHo. As in Mincy, Grossbard and Huang (2005) women having children out-of-couple (as lone mothers) may also obtain some resources from the child's father; but fathers are expected to offer more benefits to mother and child if there is cohabitation, and even more if the couple is married. The higher the BR for cohabiting and married women relative to that for women living alone the more women are likely to give birth while being in couple rather than alone (Ekert-Jaffe and Grossbard 2008).[5] Conditions in markets for cohabiting or married WiHo influence (1) the probabilities p_{co} and p_{ma} and (2) the ratio of estimated benefits to costs of being a mother in each couple form BR_{co} and BR_{ma}.

In Period 2 it is assumed that the only feasible switch is one from cohabitation to marriage, and that BR varies depending on who initiates a switch from cohabitation to marriage. The expected benefit to cost ratio of becoming a cohabiting mother in Period 1, $E(BR_{co})$, is a function of the BR of three possibilities in Period 2: staying in cohabitation (*coco*), a switch to marriage preferred by the woman (possibly in conjunction with the man), (*comaf*), and a switch to marriage preferred by the man and not in the woman's best interest, *comam*. $E(BR_{co})$ takes account of the probability of each potential switch and equals:

(2) $E(BR_{co}) = p_{coco}\, BR_{co} + p_{comaf}\, BR_{comaf} + p_{comam}\, BR_{comam}$

Replacing p_{coco} with $[(1 - p_{comaf} - p_{comam})$ and $E(BR_{co})$ in equation 1 with equation 2 we obtain that at time 0 a woman's potential child generates the following expected benefit to cost ratio:

(3) $E(BR) = p_{co}[(1 - p_{comaf} - p_{comam})\, BR_{co} + p_{comaf}\, BR_{comaf} + p_{comam}\, BR_{comam}] + p_{ma}BR_{ma} + p_a BR_a$

The more men are willing to pay for women's mothering WiHo the higher the various Benefit to Cost ratios BR and the more a woman is likely to give birth.

To obtain the predicted effect of CLM on the expected BR from becoming a mother we differentiate equation 3 according to CLM and obtain:

(4) $\partial E(BR) / \partial CLM = p_{co}\, (\partial p_{comaf}/\partial CLM)\, (BR_{comaf} - BR_{co})$
$+ p_{co}\, (\partial p_{comam}/\partial CLM)\, (BR_{comam} - BR_{co})$
$+ (\partial p_{co}/\partial CLM)\, [E(BR_{co}) - BR_{ma} - BR_a]$

According to equation 4 the predicted effect of CLM on the BR of becoming a mother, and therefore the likelihood of giving birth, is composed of three terms. The first two terms include

[5] Another law that could affect fertility via its effect on men's willingness to cohabit or marry is a law establishing unilateral divorce. Such law is likely to increase a single man's willingness to marry his unborn child's mother. In turn, a higher likelihood of marriage will make it more likely that a woman will choose marital fertility rather than its alternatives (no birth or non-marital fertility). This helps explain the finding that unilateral divorce laws led to a drop in the ratio of non-marital to marital births in the US (Drewianka 2008) and Europe (Bellido and Marcen 2014). Drewianka's explanation for this finding was similar: unilateral divorce leads to lower costs of entering marriage. All states except New York adopted unilateral divorce by 1990. Excluding NY does not substantially alter our results.

effects of CLM on the probability that a switch from cohabitation to marriage will be initiated in Period 2: these probabilities are expected to be positive since CLM laws make it easier to make such switches.

a/ The first term p_{co} $(\partial p_{comaf}/\partial CLM)$ $(BR_{comaf}$ - $BR_{co})$ is expected to be *positive*: it contains the positive effect of CLM on the probability of a transition to marriage and the difference between BR in a state that a woman prefers (comaf) and a less preferred state (co): BR_{comaf} - BR_{co} is also a positive term.

b/ The second term p_{co} $(\partial p_{comam}/\partial CLM)$ $(BR_{comam}$- $BR_{co})$ is expected to be *negative* because the partial effect of CLM on any transition from cohabitation to marriage is positive and from a woman's perspective if the man prefers a switch to marriage $(BR_{comam}$- $BR_{co})$ is negative.

c/ The third term $(\partial p_{co}/\partial CLM)$ $[E(BR_{co})$ - BR_{ma} - BR_a] includes the partial effect of CLM on p_{co}, the probability of cohabitation in the first period. This partial effect is expected to be negative to the extent that when CLM laws are applied there is potentially more to loose from cohabitation for men and women with expected incomes exceeding those of their partners. The term in square brackets is expected to be positive, so c/ is expected to be *negative*.

To the extent that terms b/ and c/ are negative and their absolute value exceeds that of positive term a/, the net effect of CLM on a woman's expected benefit to cost ratio of becoming a mother will be negative and CLM will be associated with fewer women becoming mothers. Under those assumptions it follows that:

Prediction 1. The availability of CLM reduces the likelihood that a woman gives birth.

In term c/ the effect of CLM operates via an effect on the probability of cohabitation in Period 1. In terms a/ and b/ the effect of CLM operates via an effect of CLM on the probability that cohabitation becomes marriage in the second period. Terms a/ and b/ go in opposite directions and may well cancel each other. The larger negative term c/ relative to terms a/ and b/, the more it is likely that the total effect of CLM on the probability of a birth will be negative.

Teens typically live with their parents, especially if they are younger than 18 and are not allowed to marry without parental consent. (In all the states where CLM laws changed in the period under investigation that legal age for marriage was 18.) Therefore, in all legal regimes with a minimum age at marriage requirement, and relative to the marriage probability, women younger than 18 have a higher probability of cohabitation p_{co} in Period 1 than women 18 or older. CLM in the case of these under-age women may have a larger negative effect on p_{co} and thus negative term c/ may be larger than for women 18 or older. To the extent that a rational choice model applies to all subjects, it is predicted that the total effect of CLM on the probability of a birth is more negative for women younger than 18 than for those who are older. In other words:

Prediction 2. CLM laws ruling the transition from cohabitation to marriage are more likely to have a negative effect on teenage women's likelihood of becoming a mother than on that of older women.

Women pregnant with their first child or considering conceiving their first child are more likely to live apart from their child's father than women who already have children. This is another case where there is a higher probability of choosing cohabitation in Period 1. It follows that

Prediction 3. The discouraging effect of CLM laws on the likelihood that women become first-time mothers is expected to be larger than these laws' effect on the likelihood that mothers have additional children.

The more prevalent non-marital cohabitation in a particular group, i.e. the larger p_{co} in equation 4 for members of that group, the larger the predicted effect of CLM laws on the expected benefits to cost ratio of becoming a mother. Among blacks non-marital cohabitation is more common than among whites: among all couples living together, 17% are unmarried among blacks and 11% unmarried among whites (Vespa, Lewis and Kreider 2013). Grossbard and Vernon (2014) find that relative to black women white women are twice as likely to be in couple. Therefore

Prediction 4. Relative to their impact on white women CLM laws are expected to have more impact on the likelihood that black women become mothers.

This black/white differential could also vary with how easy it is for black and white mothers to find adoptive parents for their child in case they prefer giving the child up for adoption.

Data and Sample Means

CDC Vital Statistics 1988-2012

The Vital Statistics are state level data that include the number of new births to teenage women in age groups 15-19 and 15-17 per 1,000 teenage women.[6] Table 1 reports mean state characteristics for CLM and non-CLM states. Teen fertility was higher in CLM states, 48.2 compared to 45.3 for non-CLM states. Fertility among the younger teens was correspondingly higher in CLM states as well. The population of CLM states was less educated and more religious, and received lower median household income and less generous welfare payments. The ratio of men to women among teens was slightly lower in CLM states.

Figure 1 shows a time trend in teen new births for CLM, non-CLM and transition states (i.e. states that abolished CLM during the period of observation) weighed by teen population. One can see a declining pattern in all states, although possibly at different rates. Among transition states Pennsylvania, Ohio and Idaho showed below average teen fertility rates while Georgia had above average rates.

June Fertility Supplements to the Current Population Surveys (CPS)

These micro-data are administered approximately every two years. CPS and its June extracts are large nationally representative datasets that contain information on demographic characteristics of all family members, mother's age, number of children per woman and year and month of birth of the last child. The advantage of using individual-level data is that it enables analysis by ethnicity and flexible age groups. One drawback of this supplement is that household income information is not available. We overcame this problem by including in some regressions the highest education level among household members. Another drawback of CPS is that not all

[6] Teen fertility rates for age group 15-17 are calculated by the authors.

cohabiting couples can be identified prior to 2007. Before that year only relationships between household heads and their partners were recorded while other household members were assigned either married or single status. Furthermore, before 1995 cohabiting partners were pooled with roommates. This precludes analysis of partnered relationships among teen mothers.

Our main sample includes all teenage women age 15-19, N = 51,893, with 4500-5000 records per year. We also worked with a subsample of girls age 15-17, N=31,985, and made several comparisons with samples of older women. Our variable of interest is whether the young women had a child between July of the year prior to the survey and June of the survey year. Most children born within these dates were conceived in the year prior to the survey, in year (t-1). The younger mothers' age at conception, 9-18 months prior to the survey, is therefore on average 14-16.

Around 21-25% of the sample lived in CLM states in various years, a total of 12,711 women. Sample means in Table 1 Panel B suggest that CLM states had a higher share of teen mothers. A higher percent of girls in those states were either married or had been married. Teens residing in CLM states are more likely to be Hispanic and from less educated families. Ideally, we would have liked to exclude about 5% of our sample who are foreign-born, since we do not know whether the foreign-born are as aware of CLM laws as natives. However, the information on country of birth is only available starting in 1994, which implies that had we restricted our sample to natives we would have lost data for the years 1990 and 1992. We therefore present results for a sample including foreign-born respondents, as in the CDC data. Results did not change substantially when we excluded the foreign-born.

Figure 2, Panels A and B, shows the share of girls ages 15-19 who have children and those who became mothers in the 12 months prior to the survey, by ethnicity and CLM availability. The percentage of mothers was the highest, 15.9% and 13.3%, among Hispanic and black girls in CLM states and the lowest among non-Hispanic White girls in non-CLM states, 5.7%. CLM states had a higher share of mothers among teens of all races. Panel B shows that a lower share of black girls gave birth in the last 12 months in CLM states compared to non-CLM states (5.8% versus 6.4%). In contrast, a considerably larger share of Hispanic girls became mothers in CLM states compared to non-CLM states, 9.3% vs 5.7%.

Figure 3 plots trends in motherhood among women 15 to 19 for the four transition states relative to the year that CLM was repealed. For each transition state we considered up to 7 years prior to the repeal and up to 13 years after the repeal. One can notice that the share of new mothers increased at least briefly in all transition states following the repeal. No other striking regularities are noticeable. This pattern may be a consequence of relatively small samples sizes in each year/state group (the sample size is 57-106 per year in Idaho and Georgia, and 117-248 per year in Ohio and Pennsylvania), and of the fact that data are not available every year.

Empirical Strategy

State level data

In our analysis of the state level CDC Vital Statistics data we estimated the impact of CLM on new births for state *s* in year *t* according to:

(1) $Y_{st+1} = \alpha CLM_{st} + \beta X_{st} + \delta_s + \gamma_t + u_{st}$

where Y_{st} are the logs of the number of new births to teens age 15-19 per 1000 teens, or the number of new births to teens age 15-17 per 1000 teens. Since CDC data records all births per calendar year, about one-fourth of conceptions happened in the same year and the rest occurred the year before. Our focus being on risky behavior rather than births per se, we wanted to match CLM status and other explanatory variables to the year of conception. CLM(t) thus refers to the year prior to the birth. We also experimented with estimating the model using same-year dependent variables and controls since some conceptions took place in the same calendar year as births. The results were comparable but the coefficients lost some magnitude, which makes sense since most births attributable to absence or presence of same-year CLM will not be occurring until the following calendar year.

We also estimated an alternative specification with two lead values of CLM and one lag value. Lead values CLM_{st+4} , CLM_{st+2} stand for having a CLM regime in effect in 3 or 4 years, and in 1 or 2 years. These variables allowed us to investigate whether the news of the abolition of CLM several years in the future affected the decision to engage in risky behavior today. Lag value CLM_{st-2} stands for CLM status 1 or 2 years ago; we used this variable to investigate whether a change in the law in the last two years had a delayed effect on risky behavior. Even though 2-year intervals are less precise than one year lags and leads, we wanted to make the results comparable to the results obtained using individual data where only effects of two-year lags can be estimated as data are available every other year. The second estimated equation is

$$(2)\ Y_{st+1} = \alpha 1CLM_{st} + \alpha 2CLM_{st+4} + \alpha 3CLM_{st+2} + \alpha 4CLM_{st-2} + \beta X_{st} + \delta_s + \gamma_t + u_{st}$$

The vector of state-level controls *X* includes sex ratios calculated by respondents' age and ethnicity to correct for possible influence of availability of potential partners on risky behavior. It also includes unemployment rate and log of median household income to account for economic conditions and the cost of living that may have an impact on the woman's decision to give birth. We added the share of adult population with a college degree and a measure of religiosity (share of state population who consider themselves very religious) to account for parental influence on teens' risky behavior and abortion. We added a measure of availability of abortions: share of women living in counties with no abortion provider. Finally, we included a measure of welfare generosity for single mothers, the log of maximum TANF or SNAP benefits for a family of two. Variable definitions and mean values are reported in Table 1 and explained in the notes to the table.[7]

We estimated the two models above, successively adding state-level controls, fixed time and state effects and state-specific trends. Data were weighted by the number of teenage female population. A CLM effect can be identified through cross-state variation and variation over time as four states repealed CLM over the period examined. If the availability of CLM increases fertility, all other state characteristics held constant, we will observe a positive coefficient α. If CLM reduces teen fertility, the corresponding coefficient will be negative.

δ_s are state fixed effects to account for unobservable differences in economic, legal, demographic and cultural environment that may affect individual choices, such as laws regarding child custody, state abortion policies and cultural norms;

γ_t are time dummies to capture the time trend; and

u_{ijt} are i.i.d. error terms.

[7] We have also tried including the following control variables: shares of black and Hispanic population, housing price index, the fraction of State House that is Democrat, and the number of male prisoners per 100,000 male population. None of these variables had a strong effect on the probability that the adolescent has a child.

Individual-level data

Using the individual-level CPS data, we estimated a series of models similar to (1) and (2) where the dependent variable is the probability that a woman age 15-19 had a child in the last 12 months, and the same probability for women age 15-17. In addition to the vector of state-level controls X explained above, we included several individual demographic controls, vector Z. These are age dummies and dummies for black and Hispanic ethnicity to account for differences in risky behavior among demographic groups. We included 3 dummies for the highest educational level in the household (no high school, college and graduate degree, with high school as a reference group) to correct for household income.

Ideally, we would have liked to also present separate results for blacks, Hispanics, and whites given that most adult couples – and conceivably teen couples as well - are within the same race. Indeed, there are reasons to believe that marriage market conditions differ by ethnicity in our data: means and standard deviations of the dependent variables are significantly different for the white and black subsamples. However sample sizes of non-white groups were not sufficiently large for race-specific difference-in-difference analyses. Instead, we just distinguished between whites and non-whites.

Several probit models were estimated with various sets of controls, state and year fixed effects, and with state-specific time trends. We also estimated linear probability models and the results were similar, so we choose to only report probit models.

Standard errors were clustered by state/year to adjust for correlated standard errors that are likely to arise due to common random effects at the state-year level. Clustering is necessary because the unit of observation is at the individual-level while the variation is at the state-level (Moulton 1990).

Note that CDC and CPS data on births do not fall into the same calendar slots: CDC data covers all births from January to December of year *(t+1)* according to our definition, while CPS questions ask whether a birth occurred from July to June of year *t* and *(t+1)*. This difference could possibly cause discrepancies in the effects of lag and leads estimated with the two data sets. When reconciling the results based on state data and on individual data it should be noted that CLM(t+2) in individual data corresponds roughly to CLM(t+3) in the state level data.

Results

We first report our state-level and then our individual-level results.

State-level analysis of CDC Vital Statistics

Table 2 presents OLS regression results based on state-level data on births to teens aged 15-19 per 1000 teens (Panel A) and births to teens aged 15-17 per 1000 teens (Panel B). The CDC data only report births for teenagers and all women 15 to 44. Therefore we could not compare childbearing of teens and women slightly older.

The first three regressions do not include lags and leads. Models 2 and 4 add state-level characteristics, and Models 3 and 6 add state-specific time trends The coefficients of CLM in Models 1 to 3 were not significantly different from zero for both age groups of teenagers in Panels A and B. Adding past and future CLM status (Columns 4 to 6) increased the value and significance of current year CLM in two regressions for women aged 15-19. Positive and statistically significant coefficients suggest that CLM at time *t* is associated with higher birth rates to all teens 19 and younger. The coefficients of CLM(t) are positive but not statistically significant in the younger subsample, meaning that the positive impact of CLM is stronger for older teens.

We observe significant and negative coefficients of the availability of the law 3-4 years ahead in both groups of teens. A version of results with annual leads (not presented here) points to year 3 as the year with strong negative coefficients. It appears that the knowledge that CLM will still be available in year *(t+3)* reduced births in year (*t+1)*. One explanation may be that CLM discouraged risky behavior, and the other that it encouraged abortions among teens in year *t*. Conversely, the prospect of CLM being abolished in 3 years may have been associated with higher pregnancy rates and/or fewer abortions in the current year, and therefore more births in year *(t+1)*.

When we added the effects of CLM in the current year and 1-4 years in the future we found that the combination of current and future availability of CLM was associated with fewer births for the younger group of teenagers. Conversely, for this age group, the prospect of the law being repealed was associated with more births, implying an increase in risky behavior or fewer abortions. However, when all teenagers were considered as one group (Panel A), the combined present and future impact of the repeal of CLM was close to zero.

Coefficients of other controls in Model 6 are presented in the first two columns of the Appendix. Sex ratios within age-race group had a strong impact on teen birth rates: a higher ratio of men to women increased birth rates. Lack of abortion providers had a small positive impact on teen fertility. A higher share of educated population was correlated with higher teen birth rates, all other state characteristics including trends held constant. Other state variables – median income, welfare generosity, religiosity, unemployment rate –had no significant effect on births.

Individual-level analysis of June Fertility Supplements to the Current Population Surveys (CPS)

Table 3 reports estimates of the marginal effects of CLM on the likelihood that a woman became a mother in the last twelve months. These probit marginal effects are the effects of CLM being available, holding other variables at their means. Panel A shows results for women aged 15-19, Panel B includes only younger women aged 15-17, and Panel C is for women aged 20-22. Models in Columns 1-3 included only CLM and Models 4-6 included one lag value and two lead values of CLM. All models included individual characteristics and state and year fixed effects, Models 2 and 4 added state-level characteristics, and Models 3 and 6 added state-specific time trends. The full list of controls can be found in the Appendix showing full estimates of Model 6 for women aged 15 to 17 and 15 to 19.

The first three regressions show negative coefficients of CLM, statistically significant at 5% in the regressions for all teens and at 1% in the regressions for the younger teens. It suggests that adolescents have lower probabilities of risky behavior when their state has CLM at time *t.* However, CLM does not have a significant coefficient in regressions 1 to 3 in Panel C reporting regressions for women aged 20 to 22.

When leads and a lag of CLM were added (Columns 4 to 6), CLM(t) is no longer statistically different from zero in the case of teens but it is in the case of women ages 20 to 22, and coefficients are negative. For young teens aged 15 to 17 anticipation of CLM within two years discouraged births. In the case of women in their early twenties (Panel C) the opposite is the case: anticipation of CLM in the next two years raised the probability of a birth. The total effect of CLM on women aged 20 to 22 on probability of birth was not clearly negative.

The marginal effect of CLM on births is thus very sensitive to the age at which women were observed. The results for teens based on state level and individual level data are compatible; all significant effects of CLM are negative and the discrepancy in the coefficients reported in Tables 2 and 3 can be explained as a function of whether annual data are available or not. That teen births responded more negatively to CLM than births to women in their early twenties is compatible with our conceptual framework.

Table 4 reports further results for women aged 15 to 19 who are either non-Hispanic white (Panel A) or black or Hispanic (Panel B). When only CLM at time t is included in the regressions, and state-specific trends are not taken into account, CLM was associated negatively with births to non-white teenaged women. In contrast, this was not the case for white women whose fertility does not seem to respond to CLM at time t. Not surprisingly, the coefficients in the first two columns of Panel B in Table 4 are larger (in absolute terms) than those in the corresponding models in Panel A in Table 3. It thus seems that our findings for the entire sample are dominated by CLM effects for blacks and Hispanics. This is consistent with Prediction 4 and could be related with the higher proportions of unmarried couples in the latter groups.

Robustness checks

Our results are robust to various changes in specification. We experimented with excluding New Hampshire, where CLM is available only in case of a partner's death, and Nebraska, where marriage age is 19 and not age 18 as in other states. We also ran regressions in which we excluded foreign born women in 1994 and later years. Yet our main results were not sensitive to these changes.

We also repeated our CPS analysis using a different definition of motherhood and a different subsample. We estimated the probability that a woman had her first child in the last 12 months using a subsample of women who never had a child before. As reported in Table 5, Panels A and B, the coefficients of CLM are very similar to those in Table 3. This implies that we do not have evidence leading us to either confirm or reject Prediction 3. We also estimated the same models using data from four transition states only. As shown in Panel C the coefficients have the same sign but overall are less significant than those obtained from a larger sample.

The difference-in-difference (DD) approach that we used for identification has been shown to suffer from a serial correlation problem. As a result, the standard errors of DD estimators are often underestimated and thus the statistical significance of the coefficients is overestimated. Following Bertrand, Duflo and Mullainathan (2004), we attempted to deal with possible serial correlation of errors by estimating a panel data model using individual-level data aggregated by year/state cells. We computed these estimates and recorded the results in Table 6. First, we regressed the binary birth data on personal characteristics. Then we calculated means of residuals by year/state and regressed the mean of residuals on CLM, state characteristics, and state and year fixed effects. The coefficients of the linear probability models were of similar sign but not as strong as the individual-level results.

Finally, we also conduct a falsification test using another teenage behavioral outcome not expected to be affected by CLM. If CLM indeed does not affect this behavior in our analyses, then we can be more confident in the results tying teen fertility to CLM. Other risky behaviors, such as smoking, alcohol use or drug use are potential candidates for such a test. We choose to estimate models for teen tobacco smoking. Data on smoking behavior among respondents age

15-19 were drawn from Current Population Survey Tobacco Supplements for 1992-2007[8]. Later years include only respondents age 18+, hence they are excluded from the analysis. Respondents with missing information on smoking were excluded from the sample. Means and standard deviations for all variables in this dataset – such as the share of teens living in CLM states and the distribution of respondents by age and race – are similar to the sample characteristics of the Fertility Supplement described above.

Table 7 reports results that indicate no consistent effect of CLM on teen tobacco use. Other variables controlled for in the regressions include dummies for gender, age, race, highest education level in the household, teen's student status, presence of own children, and a dummy for whether teen's mother or father smokes. State characteristics are the same as in regressions for teen birth rates. One should keep in mind that all teen risky behaviors may be positively correlated via unobservable individual characteristics or unobserved outcomes: for instance, a teen who drops out of school to give birth may be more likely to socialize with other dropouts and experience peer pressure for other risky behaviors. Nevertheless, the absence of results for teenage smoking raises our confidence in our findings regarding CLM and teen fertility.

Conclusions

We analyzed whether the likelihood that teenage women engage in risky behavior, as measured by teen birth rates and the likelihood that they became mothers over the last 12 months, is linked to whether Common Law Marriage (CLM) was available or expected to be available in the near future in the state of residence at the time of pregnancy. Based on a comparison of the states that abolished CLM and those that did not we found that the abolition of CLM had a small positive effect on births to teen mothers. Availability of CLM three years in the future had a larger impact than its availability around the time of potential conception. Younger teens aged 15 to 17, or their partners, were more responsive to the availability of CLM than women aged 15 to 19 or their partners. For women aged 20 to 22 we did not find a net effect of CLM on births, suggesting that they or their partners did not change their risky behavior as a function of availability of CLM. Our findings are based on analyses of both state-level CDC Vital Statistics data and individual-level information from the CPS. The effect of CLM was small and the fact that some states repealed CLM was not sufficient to reverse the overall trend towards lower teen fertility observed in the US in recent years.

That CLM availability was associated negatively with teen births is consistent with the rational choice model that we presented. Women or their partners may be more careful regarding birth control or more willing to opt for an abortion if they prefer to avoid marriage and CLM availability may raise the threat of an unwanted marriage imposed by a partner relying on CLM. Teenagers engaged in risky behavior are considerably more likely to do so outside marriage than their older counterparts, given that they often are too young to marry. Consequently the potential threat presented by CLM is more relevant to male and female teens than to women in their twenties or their partners. Likewise, given that 18 is the legal age for marriage in most states, relative to their older counterparts teens younger than 18 are more likely to be in situations of cohabitation potentially leading to unwanted unilateral marriage. Our findings of a stronger negative effect of CLM for teenagers than for older women and for teens under 18 than for all teens combined thus support the rational choice framework presented here.

[8] IPUMS-CPS, University of Minnesota, www.ipums.org.

Our analyses also indicated that it is not solely the availability of CLM at the time of conception that matters, but also the prospect of the law still being in place in the near future.

The last time a US state repealed CLM was when Pennsylvania repealed it in 2005. The eleven states that still have CLM may want to take our study into account. Its policy implication is that if teenage births ought to be minimized it is preferable not to repeal CLM for that may cause a small rise in teen births. At the same time, this exploratory study's results are not large enough to provide a justification for the other states to do what Utah did in 1987 and reinstate CLM.

In an earlier study based on a similar method of analysis we found that CLM discouraged couple formation by men with more than a high school education (Grossbard and Vernon 2014). Relative to less educated men these men stand to lose more from an unwanted transition from cohabitation to marriage initiated by their partner. Following the same logic it is possible that relative to less talented counterparts, when CLM is available young men with better educational and career prospects are more reluctant to engage in risky behavior that could lead to an unwanted marriage. This could not be tested as most teens do not attend college and we did not have information on teens' school performance. However, if such associations between CLM, education, and risky behavior were to be observed, one may find that the availability of CLM encourages accumulation of human capital via added schooling or on-the-job training, especially for the men who potentially father the children of teenagers. They may have more incentives to use contraceptives when CLM is available and the threat of a future unwanted marriage is more of a reality.

A broader implication of our study is that rational choice models seem to apply to teens and to a risky behavior such as engaging in unprotected sex. Otherwise, it is hard to explain why teens responded to availability of CLM at the time of potential conception or a few years later.

Given that this is a first study of its kind, our policy recommendations should be interpreted as suggestions. A further reservation is that reinstating CLM where it was abolished may not be desirable in light of other findings: the law appears to have delayed gender convergence in labor supply (Grossbard and Vernon 2015) and to have discouraged couple formation among women at the conventional age of childbearing (ages 18 to 35) according to Grossbard and Vernon (2014).

Further research is needed on this topic and on related risky behaviors by teens, such as heavy drinking. A major limitation of this study was lack of information on kind of marriage when observing a married couple in a CLM state. It is hoped that new data sets will include information on whether couples are married in a conventional sense or as a result of CLM in those states where the law is available. Such data will allow researchers to better assess the impact of the law on births and other outcomes. For example, data from other countries or regions practicing CLM could be analyzed. We also hope that there will be further investigations of the effects of CLM on various subgroups of the population categorized by ethnicity or immigrant status.

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Table 1. Means of state-level characteristics.

Variable	non-CLM	CLM
A. State-level characteristics:		
Births to women age 15-17 per 1000 of women age 15-19	24.6	28.2
Births to women age 15-19 per 1000 of women age 15-19	45.3	48.2
Unemployment rate#	5.6	5.4
Median household income in 2010 dollars	50,372	48,648
Welfare benefits, in 2010 dollars	738	697
Sex ratio	1.06	1.05
Share of women 15-44 in counties with no abortion provider#	41.4	42.2
Number of women age 15-44, millions	1.96	1.48
Number of women age 15-19, millions	0.31	0.24
Share of religious population	39	42
Share of adults age 25+ with college degree	24.2	23.9
State population, millions	5.86	4.51
Observations	937	338
B. Individual-level characteristics:, women age 15-19		
Ever had a child	7.5	9.2
Had a child in the past 12 months#	4.1	5.1
Had first child in the past 12 months#	2.6	3.3
Currently married	2.3	3.2
Ever married	3.2	4.2
Black#	16.0	16.2
Hispanic	13.8	19.0
Highest educ in family: No high school	8.5	10.4
Highest educ in family: College#	18.7	17.1
Highest educ in family: Graduate#	12.3	10.8
Observations	39,182	12,711

the difference is not statistically significant at 5% level.

Sources of information for state-level data:

Vital statistics data on births is from CDC http://www.cdc.gov/nchs/data_access/vitalstats/VitalStats_Births.htm

Median household income is in Table H-8 at http://www.census.gov/hhes/www/income/data/historical/household/. Unemployment rates are annual averages by state obtained from BLS (http://www.bls.gov/data/).

Sex ratios are calculated from Census state population estimates by dividing the number of men in each 5-year age group by the number of women who are 2 years younger. For example, in order to get a sex ratio for women aged 18-22, we divide the number of men aged 20-24 by the number of women aged 18-22. Sex ratios are calculated separately for white and black population and for the total population. 'Other race' is assigned sex ratios for total population.

Population data by race and age are obtained from Census:
http://www.census.gov/popest/data/intercensal/state/state2010.html and
https://www.census.gov/popest/data/state/asrh/2012/SC-EST2012-ALLDATA6.html

Welfare benefits are maximum TANF+SNAP benefits for a family of two, in 2010 dollars obtained from the University of Kentucky Center for Poverty Research http://www.ukcpr.org/AvailableData.aspx

Share of adults age 25+ with college degree are from Census: http://www.census.gov/hhes/socdemo/education/

All state-level mean values other than state populations are weighted by state teenage women population.

Share of religious population is from U.S. State Religion & Society Data, http://www.gallup.com/poll/125066/State-States.aspx , it refers to a share of state population who consider themselves 'very religious'. The values are not available for some years, they are imputed as average values between years for which the values are known.

Data on abortion providers are from Guttmacher Institute http://www.guttmacher.org/datacenter/

Individual characteristics are from June CPS 1990-2010. Individual-level mean values are weighted using survey weights.

Table 2. Birth rates, CDC state-level data, 1988-2012.

	1	2	3	4	5	6
Panel A. Log birth rates among women age 15-19, N=1,224						
CLM(t+4)				**-0.03**	**-0.037**	-0.025
				[0.013]*	**[0.017]***	[0.014] †
CLM (t+2)				-0.01	-0.004	-0.006
				[0.008]	[0.010]	[0.010]
CLM (t)	-0.044	-0.014	-0.007	0.007	**0.025**	**0.018**
	[0.029]	[0.021]	[0.021]	[0.006]	**[0.005]****	**[0.004]****
CLM(t-2)				-0.03	-0.019	-0.026
				[0.022]	[0.013]	[0.021]
Panel B. Log birth rates among women age 15-17, N=1,173						
CLM(t+4)				**-0.061**	**-0.073**	**-0.045**
				[0.018]**	**[0.024]****	**[0.018]***
CLM (t+2)				-0.004	-0.001	0.001
				[0.011]	[0.015]	[0.009]
CLM (t)	-0.043	-0.026	0.001	0.011	0.028	0.022
	[0.039]	[0.037]	[0.027]	[0.009]	[0.014] †	[0.014]
CLM(t-2)				-0.021	-0.018	-0.027
				[0.029]	[0.025]	[0.031]
State characteristics		X	X		X	X
Year & state fixed effects	X	X	X	X	X	X
State-specific trends			X			X

Notes. Births to teens are per 1000 female population age 15-19. Robust standard errors in brackets. R-square ranges between 0.86-0.99.

Here and in the rest of the tables: † $p < .10$, * $p < .05$, ** $p < .01$, *** $p < .001$. Statistically significant results are in bold.

Table 3. Probit marginal effects of CLM on the probability that a woman had a child in the last 12 months, CPS 1990-2010 individual level data. All races, by age group.

	1	2	3	4	5	6
Panel A. Women age 15-19, N=51,893						
CLM(t+4)				0.019	0.019	0.014
				[0.015]	[0.015]	[0.017]
CLM (t+2)				**-0.027**	**-0.028**	**-0.026**
				[0.008]**	**[0.008]****	**[0.010]****
CLM (t)	**-0.010**	**-0.012**	**-0.017**	0.006	0.007	0.005
	[0.005]*	**[0.005]***	**[0.008]***	[0.005]	[0.006]	[0.007]
CLM(t-2)				0.006	0.004	0.006
				[0.005]	[0.005]	[0.007]
Panel B. Women age 15-17, N=31,985						
CLM(t+4)				0.004	0.005	0.009
				[0.010]	[0.010]	[0.010]
CLM (t+2)				**-0.021**	**-0.022**	**-0.028**
				[0.007]**	**[0.006]****	**[0.007]****
CLM (t)	**-0.018**	**-0.019**	**-0.020**	0	0	0.012
	[0.004]**	**[0.004]****	**[0.004]****	[0.010]	[0.010]	[0.015]
CLM(t-2)				0.004	0.002	0
				[0.007]	[0.008]	[0.007]
Panel C. Women age 20-22, N=28,637						
CLM(t+4)				-0.041	-0.043	-0.05
				[0.024] †	[0.024] †	[0.028] †
CLM (t+2)				**0.106**	**0.104**	0.079
				[0.045]*	**[0.043]***	[0.052]
CLM (t)	0.014	0.015	0.009	**-0.044**	**-0.042**	**-0.036**
	[0.020]	[0.022]	[0.038]	**[0.011]****	**[0.011]****	**[0.012]****
CLM(t-2)				0.015	0.016	0.004
				[0.014]	[0.014]	[0.018]
State characteristics		X	X		X	X
Year & state fixed effects	X	X	X	X	X	X
State-specific trends			X			X

Notes: All regressions include demographic controls. Robust standard errors in brackets. Standard errors are clustered by state-year.

Table 4. Probit marginal effects of CLM on the probability that a woman had a child in the last 12 months, CPS 1990-2010 individual level data, by race.

	1	2	3	4	5	6
Panel A. White women age 15-19, N=36,019						
CLM(t+4)				0.027	0.025	0.008
				[0.016]+	[0.016]	[0.015]
CLM (t+2)				**-0.028**	**-0.024**	**-0.025**
				[0.006]* **	**[0.007]*** **	**[0.008]** **
CLM (t)	-0.002	-0.003	-0.011	**0.024**	0.016	**0.018**
	[0.006]	[0.006]	[0.008]	**[0.012]***	[0.010]+	**[0.008]***
CLM(t-2)				-0.001	-0.002	0
				[0.008]	[0.007]	[0.007]
Panel B. Non-white women age 15-19, N=15,836						
CLM(t+4)				-0.005	-0.006	0.013
				[0.016]	[0.016]	[0.021]
CLM (t+2)				0.016	0.007	0.011
				[0.023]	[0.022]	[0.034]
CLM (t)	**-0.025**	**-0.031**	-0.024	**-0.043**	**-0.039**	**-0.037**
	[0.009] **	**[0.009]*** **	[0.015]	**[0.010]*** **	**[0.011]*** **	**[0.016]***
CLM(t-2)				0.023	0.017	0.019
				[0.015]	[0.014]	[0.026]
State characteristics		X	X		X	X
Year & state fixed effects	X	X	X	X	X	X
State-specific trends			X			X

Table 5. Robustness check 1. Probit marginal effects, first time mothers. CPS.

	1	2	3	4	5	6
Panel A. Women age 15-19, N=49,479						
CLM(t+4)				0.009	0.009	0.004
				[0.005] †	[0.006]	[0.005]
CLM (t+2)				**-0.015**	**-0.016**	**-0.016**
				[0.003]**	**[0.003]****	**[0.004]****
CLM (t)	**-0.007**	**-0.007**	**-0.008**	0.009	0.009	0.012
	[0.003]*	**[0.003]***	**[0.004]***	[0.005] †	[0.005] †	[0.006] †
CLM(t-2)				-0.003	-0.003	0
				[0.003]	[0.003]	[0.004]
Panel B. Women age 15-17, N=30,710						
CLM(t+4)				-0.002	-0.003	-0.002
				[0.009]	[0.009]	[0.009]
CLM (t+2)				**-0.012**	**-0.013**	**-0.017**
				[0.005]*	**[0.005]***	**[0.005]****
CLM (t)	**-0.012**	**-0.013**	**-0.012**	0.003	0.005	0.019
	[0.003]**	**[0.003]****	**[0.003]****	[0.007]	[0.008]	[0.011] †
CLM(t-2)				-0.002	-0.003	-0.003
				[0.005]	[0.004]	[0.004]
Panel C. Women age 15-19, transition states only, n=5,410						
CLM(t+4)				0.017	0.015	0.017
				[0.016]	[0.016]	[0.016]
CLM (t+2)				**-0.029**	-0.021	-0.028
				[0.014]*	[0.014]	[0.016] †
CLM (t)	-0.01	-0.009	-0.009	0.014	0.008	0.015
	[0.006] †	[0.006]	[0.007]	[0.017]	[0.016]	[0.019]
CLM(t-2)				-0.001	0.008	0.011
				[0.005]	[0.007]	[0.008]
State characteristics		X	X		X	X
Year & state fixed effects	X	X	X	X	X	X
State-specific trends			X			X

Notes: All regressions include demographic controls. Robust standard errors in brackets. Standard errors are clustered by state-year. Women who had a child prior to 12 month before the survey are removed from regressions for first time mothers; this does not alter results. Only 589 girls in this age group are married; excluding them does not alter the results.

Table 6. Robustness check 2: OLS estimates of linear probability models with data aggregated by state/year.

	Age 15-19, N=561				Age 15-17, N=561			
	1	2	3	4	5	6	7	8
CLM(t+4)			0.007	0.006			0.002	-0.012
			[0.015]	[0.017]			[0.015]	[0.017]
CLM (t+2)			-0.025	-0.024			-0.017	-0.028
			[0.030]	[0.032]			[0.029]	[0.031]
CLM (t)	-0.012	-0.015	-0.002	-0.004	-0.014	-0.027	-0.014	-0.007
	[0.009]	[0.015]	[0.030]	[0.032]	[0.009]	[0.014] †	[0.030]	[0.031]
CLM(t-2)			0.008	0.009			0.016	0.011
			[0.014]	[0.017]			[0.014]	[0.017]
State characteristics	X	X	X	X	X	X	X	X
Year & state fixed effects	X	X	X	X	X	X	X	X
State-specific trends		X		X		X		X
R-squared	0.23	0.3	0.25	0.31	0.22	0.3	0.24	0.31

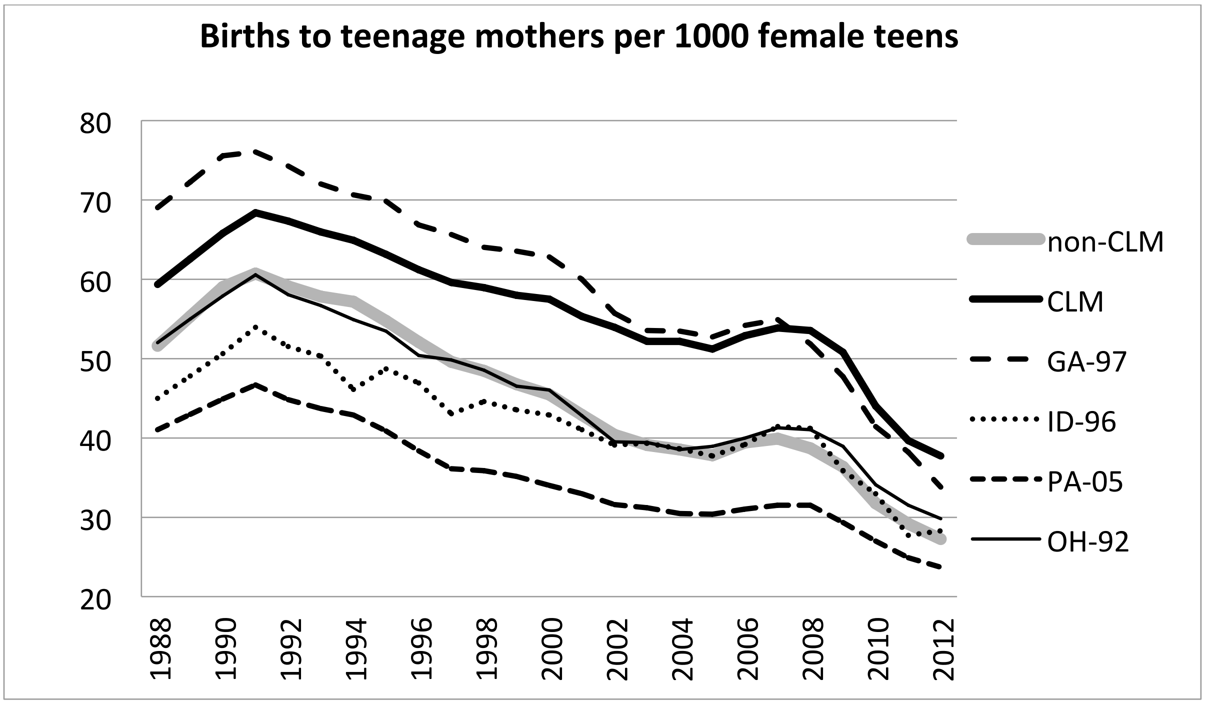
Notes: To obtain these estimates, we first regress data on individual-level controls and collect residuals. Then we compute average residuals by year and state, and regress them on state characteristics and year and state fixed effect with and without state time trends.

Table 7. Falsification test. Probit marginal effects of CLM on the probability that a teen is a smoker. CPS Tobacco Supplements 1992-2007.

	1	2	3	4	5	6
Boys and girls age 15-19, N=138,601						
CLM(t+4)				0	0	0.009
				[0.002]	[0.001]	[0.005]+
CLM (t+2)				-0.009	-0.006	-0.007
				[0.019]	[0.019]	[0.017]
CLM (t)	-0.003	-0.003	0.002	0.009	0.005	0.01
	[0.009]	[0.009]	[0.002]	[0.010]	[0.009]	[0.012]
CLM(t-2)				-0.008	-0.006	**-0.01**
				[0.005]	[0.005]	**[0.003]** *
State characteristics		X	X		X	X
Year & state fixed effects	X	X	X	X	X	X
State-specific trends			X			X

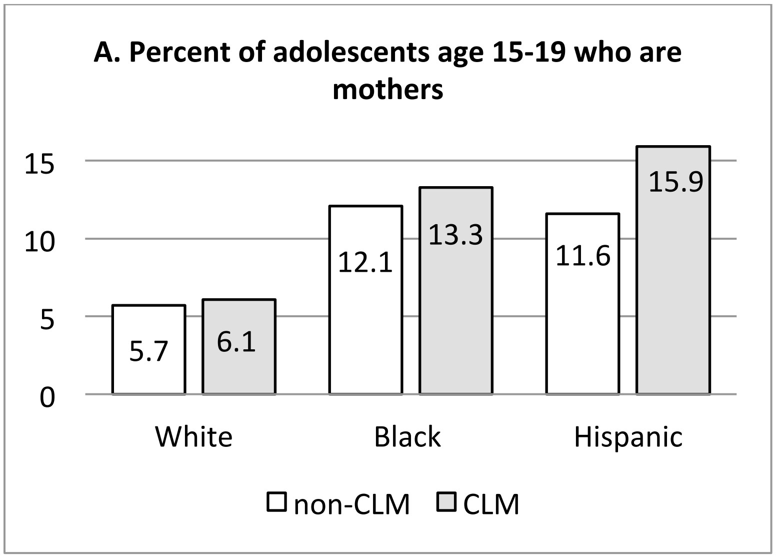
Notes: The dependent variable is a binary variable with value 1 for smokers. A smoker is defined as a respondent who answered yes to the following two questions 'Have you smoked 100 or more cigarettes?' and 'Do you currently smoke on either some days or every day?' Tobacco Supplements for 1992-2007 contain information on teens aged 15-19.

Figure 1. Births to teenage mothers age 15-19 per 1000 female teens. CDC state-level data, 1988-2012.



Note. The graph shows birth rates to mothers age 15-19. Data are weighted by the corresponding state female teen population age 15-19.

Figure 2. Percent of adolescent girls age 15-19 who ever gave birth and those who gave birth for the first time. June CPS 1990-2010.



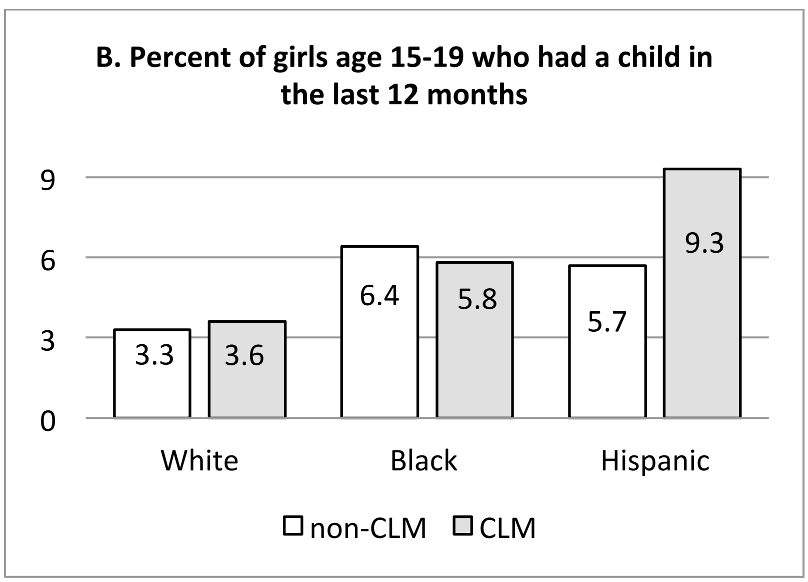
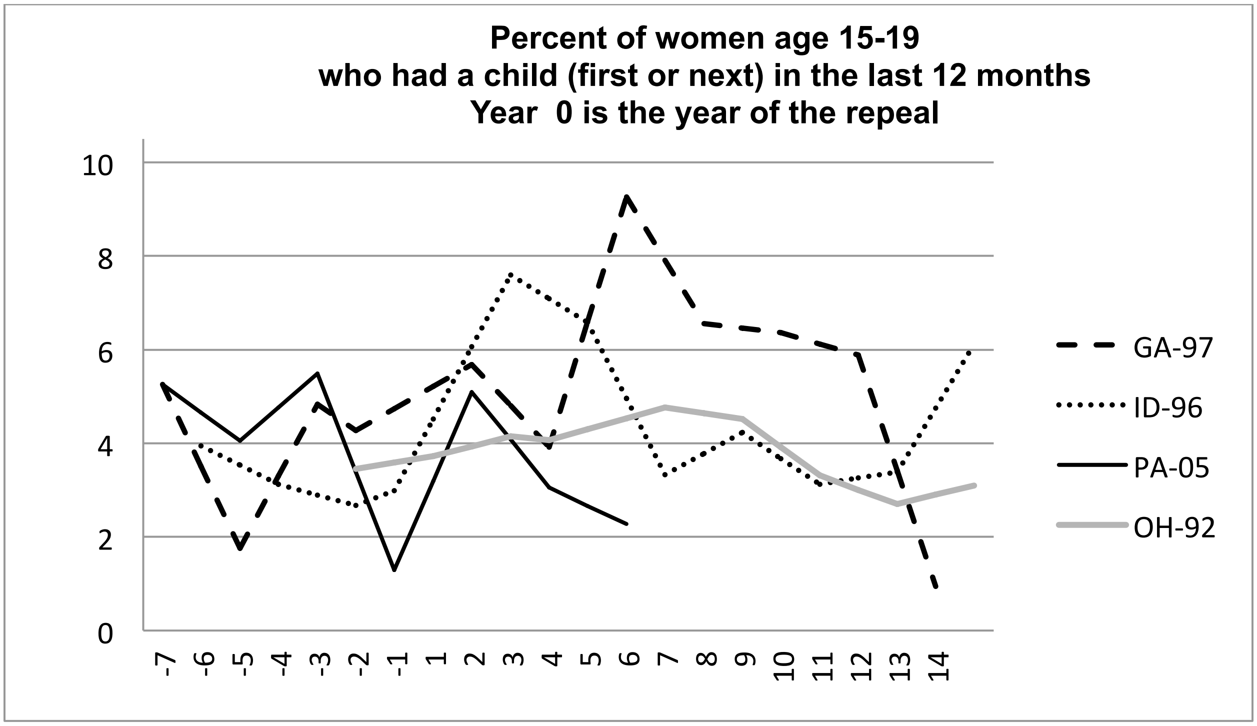


Figure 3. Evolution of motherhood rates for women age 15-19, June CPS individual-level data 1990-2010. Transition states.



Note: The graph shows percent of women age 15-19 who had a child (first or next) in the last 12 months. Year 0 is the year of the CLM repeal. The data are weighted using survey weights.

Appendix. OLS coefficients in CDC regressions. Marginal effects in CPS regressions.

	CDC (Table 2, model 6)		CPS (Table 3, model 6)	
	Age 15-19	Age 15-17	Age 15-19	Age 15-17
CLM(t+4)	-0.025	-0.045	0.014	0.009
	[0.014] †	[0.018]*	[0.017]	[0.010]
CLM (t+2)	-0.006	0.001	-0.026	-0.028
	[0.010]	[0.009]	[0.010]**	[0.007]**
CLM (t)	0.018	0.022	0.005	0.012
	[0.004]**	[0.014]	[0.007]	[0.015]
CLM(t-2)	-0.026	-0.027	0.006	0
	[0.021]	[0.031]	[0.007]	[0.007]
Sex ratio, race-specific	2.499	2.982	-0.018	0.005
	[0.593]**	[0.863]**	[0.085]	[0.108]
Welfare benefits	0.084	0.104	0.02	-0.001
	[0.048] †	[0.065]	[0.019]	[0.018]
Unemployment rate	0.007	0.003	-0.002	0
	[0.006]	[0.006]	[0.002]	[0.002]
Share of adults with college degree	0.004	0.006	0	0.001
	[0.002] †	[0.002]*	[0.001]	[0.001]
Log of median state income	0.028	0.015	0.002	-0.009
	[0.050]	[0.051]	[0.028]	[0.027]
No abortion provider	0.002	0.003	0	0
	[0.001]*	[0.001]*	[0.000]	[0.001]
Share of religious population	0.006	0.005	0.002	0.004
	[0.004]	[0.005]	[0.003]	[0.003]
Black			0.017	0.012
			[0.005]**	[0.006] †
Hispanic			0.017	0.007
			[0.004]**	[0.004] †
Max educ in family: no high school			0.037	0.017
			[0.005]**	[0.005]**
Max educ in family: college			-0.013	-0.002
			[0.002]**	[0.003]
Max educ in family: graduate degree			-0.016	-0.003
			[0.003]**	[0.003]
Age 16			0.019	0.014
			[0.005]**	[0.004]**
Age 17			0.036	0.026
			[0.006]**	[0.004]**
Age 18			0.053	
			[0.007]**	
Age 19			0.072	
			[0.007]**	
Year and state fixed effects	X	X	X	X
State-specific time trends	X	X	X	X
Observations	1224	1173	51893	31985